International Risk-Sharing: Macroeconomic and Financial Aspects

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International Risk-Sharing: Macroeconomic and Financial Aspects

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(со резиме на македонски јазик)

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Metodij Hadzi-Vaskov

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PROMOTOR: Prof. dr. C.J.M. Kool

НА МАМА И ТАТО

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Metodij Hadzi-Vaskov
Utrecht, May 2008

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Chapter 1

Introduction

1.1 Motivation and Background

The process of international financial integration, which accelerated in the past decades, can potentially bring numerous benefits to the world economy.[1] The opportunity for international risk-sharing is one of the most important, and at the same time, one of the least controversial potential welfare benefits from this process.[2] As long as different national economies are not perfectly linked, there are possibilities for risk-sharing through cross-border trade in assets and goods. By getting the opportunity for pooling risks in a larger, global market, the residents in these economies can diversify the idiosyncratic, country-specific portion of the risk they face, and thereby, obtain additional welfare gains that are not available within their national borders. In turn, the condition for perfect (optimal) international risk-sharing suggests complete diversification of all idiosyncratic (country-specific) risks, so that each country faces only (the proportional share of) the aggregate (global) systemic risk.[3]

Inevitably, the future advancement of the process of financial globalization is conditional on the actual welfare benefits that it delivered in the past. How-

[1] A general indicator for this acceleration is the dramatic increase in the sum(s) of gross external assets and liabilities relative to GDP in this period, see Lane and Milesi-Ferretti (2001, 2003, 2006).

[2] See Kose et al. (2007) or IMF (2007). For other potential benefits see Obstfeld (1998), Prasad et al. (2003), Kose et al. (2007), and IMF (2007).

[3] For microeconomic tests on the condition of perfect risk-sharing see Cochrane (1991) or Mace (1991).

ever, alternative approaches encountered in the academic literature suggest very different results with respect to the degree of international risk-sharing that actually takes place.[4] Therefore, in light of the anticipated inclusion of additional (developing) countries in the financial integration process, it is important to further analyze these alternative methods in an attempt to find a conclusive framework for quantifying the degree of international risk-sharing.

1.2 Aim of the Thesis

This thesis focuses on the macroeconomic and financial dimensions (aspects) of measuring international risk-sharing. First, it aims to investigate the extent to which risk-sharing takes place across countries using macroeconomic and financial approaches. Second, it attempts to present the evidence in a compact/synthetic frame. As it is central to both of these approaches, and naturally "links" them, the (real) exchange rate takes the role of conductor in this synthesis.

1.3 Place in the Literature

This thesis is related to several different, though occasionally intertwined strands in the macroeconomic and financial literature. First, the core issue investigated in this thesis is the measurement of international risk-sharing. In this way, it is related to several alternative perspectives for quantifying the level of cross-country risks diversification. This "*core*" part of the literature related to the research undertaken in the thesis encompasses the following three strands:

- macroeconomic models of the real business cycle (RBC) type based on a specific form of the utility function - usually power utility of the constant

[4]Macroeconomic evidence suggests poor risk-sharing: consumption growth rates across OECD countries are less correlated than output growth rates (Backus et al. 1992), consumption volatility increased relative to output volatility for many emerging economies recently (Kose et al., 2003), and relative consumption is typically negatively related to real exchange rate changes (Backus and Smith, 1993). On the other hand, asset-markets-based measures suggest very high degrees of risk-sharing (Brandt et al., 2006).

relative risk-aversion (CRRA) class. Empirical tests of the (perfect) risk-sharing condition in these models are usually based on macroeconomic data series: aggregate consumption and aggregate output/income (Backus et al. 1992; Backus and Smith, 1993; Obstfeld, 1994, 1995; Lewis, 1996, 1999; Pakko, 1998, 2004; Van Wincoop, 1994, 1999) [point of reference for this thesis, chapters 2, 6, and 7];

- measures based on portfolio theory (French and Poterba, 1991; Tesar and Werner, 1998; Ahearne, Griever, and Warnock, 2004; Baxter and Jermann, 1997)[5][chapters 2 and 4];

- international asset pricing framework grounded on non-arbitrage conditions (Backus, Foresi, and Telmer, 1996; Backus, Foresi, and Telmer, 2001; and Brandt, Cochrane, and Santa-Clara, 2006) [6][chapters 4 and 5].

Additional to this "*core*" part of the literature, the investigation in some chapters is related to one, or more "*lateral*" literature strands.

- alternative channels of international risk-sharing: portfolio diversification (Lewis, 1999, 2000; Sorensen et al., 2007), fiscal policy (Asdrubali, Sorensen, and Yosha, 1996; Sorensen and Yosha, 1998), FDI (Albuquerque, R., 2003), investment in multinationals (Heston and Roewenhorst, 1994; Errunza et al., 1999; or Rowland and Tesar, 2004), current and capital account fluctuations (Sachs, 1982; Baxter, 1995; Glick and Rogoff, 1995; or Obstfeld and Rogoff, 1995)[chapters 2 and 4];

- tests related to international (equity return or interest rate) parity conditions (Chinn and Meredith, 2005; Chinn, 2006; Cappiello and De Santis, 2005, 2007, Lustig and Verdelhan, 2007)[chapters 3, 4, 6, and 7];

- disconnect of real exchange rates from fundamentals and the (non)neutrality of the nominal exchange rate regimes (Mussa, 1986; Baxter and Stockman, 1989; Dixon, 1999; Flood and Rose, 1999; Lothian and McCarthy, 2002; MacDonald, 1999)[chapters 5, 6, and 7].

[5]For an application of portfolio theory with macroeconomic data, see Imbs and Mauro, 2007.

[6]For a general overview of asset pricing models, see Cochrane, 2004.

1.4 Thesis Outline

This introductory chapter closes with an outline of the main issues discussed in the remaining 7 chapters of this thesis.

Chapter 2 deals with alternative channels for international risk-sharing. It starts by briefly reviewing some important studies on the macroeconomic and financial aspects of international risk-sharing. This review of two interrelated strands of the literature (macroeconomic and financial) brings forward the cross-border diversification of portfolio holdings as the most important driving force behind the process of international risk-sharing. While acknowledging the preponderant role of international portfolio diversification in the case of advanced, industrial countries, the discussion is more cautious when it comes to developing countries. It identifies workers' remittance flows as an alternative channel through which the process of cross-border macroeconomic risk-sharing might take place. This is motivated by three well-documented characteristics of the workers' remittance flows to developing countries: first, these flows are the most stable source of external finance and foreign exchange for many developing countries; second, they do not display high levels of procyclicality; and third, they jump sharply after economic (financial) crises hit a particular developing country. In sum, these three characteristics make them especially suitable instruments for consumption smoothing in the face of idiosyncratic country-specific shocks.

The empirical part of this chapter investigates the importance of workers' remittances for improving international risk-sharing. At the beginning, it presents some stylized facts about the degree of international risk-sharing for industrial and developing countries. Afterwards, it focuses on the question whether developing countries with above-average levels of workers' remittances inflows per capita also achieve above-average levels of international risk-sharing in (aggregate) consumption. The analysis looks at the group of all developing countries as well as specific sub-groups: less financially-integrated countries (LFIs), more financially-integrated (MFIs), and transition economies.

Chapter 3 elaborates on the possibility that the empirical procedure conventionally used to test for the uncovered interest rate parity (UIP) condition might be biased. In the context of the introduction to this section, testing for (covered or uncovered) interest rate parity across countries is one of the most frequently used price-based measures of (the intensity of) the process of international financial integration (financial globalization). This condition suggests that the expected returns on two comparable (risk-free) assets denominated in different currencies should be equal, if these assets only differ with respect to the currency of denomination. Though a theoretically sound concept, UIP is consistently rejected in empirical tests. This chapter begins with a demonstration that the slope coefficient estimator from the empirical procedure conventionally used to test for UIP may be biased if the underlying data-generating process slightly differs from the theoretically expected one. The main aim of the chapter is to investigate the extent to which this estimation bias is important for the slope coefficient estimates in conventional tests of UIP. It is important to note that the analysis in this chapter does not attempt to "rehabilitate" the UIP condition, nor does it attempt to completely resolve the UIP puzzle.

Chapters 4 and 5 focus on the measurement of international risk-sharing based on the stochastic discount factor (SDF) approach. Contrary to most evidence from macroeconomic studies based on international real business cycle models and from financial studies rooted in portfolio theory, this approach suggests a high degree of cross-country diversification in macroeconomic risks. Starting with the basic (international) asset pricing condition, which states that real exchange rate changes (should) exactly equal the difference between the stochastic discount factor movements for the corresponding two countries, the (bilateral) SDF approach shows that: first, there is a lot of macroeconomic risk to be shared across countries because the discount factors are very volatile; and second, a very large portion of this macroeconomic risk is actually shared across countries because the real exchange rates are much less volatile. Chapters 4 and 5 present some limitations and extensions of this approach.

Chapter 4 presents the theoretical framework for the (bilateral) SDF approach to international risk-sharing and replicates some of the results found earlier in the literature. Moreover, it points out two main limitations (or inherent incoherencies) of this bilateral setting. As an attempt to account for these incoherencies, it presents a three-country extension of the SDF model and calculates risk-sharing figures for this trilateral setting.

As a follow-up to chapter 4, chapter 5 emphasizes the importance of the nominal exchange rate behavior (regime) for the SDF-based measures of international risk-sharing. For this purpose, it presents evidence about two episodes of rigid/fixed nominal exchange rate: the Eurozone countries before and after the introduction of the common currency (time period 1993-2005), and several emerging market economies with a fixed or very rigid nominal exchange rate against the US dollar over the same time period.

Chapter 6 investigates the importance of the nominal exchange rate fluctuations for the anomalous correlation between the relative consumption growth and the real exchange rate changes. This anomalous relationship runs counter to the condition for efficient international (consumption) risk-sharing derived in (international) real business cycle models, which suggests that consumption should be relatively higher in countries that experience a drop in relative prices. Thereby, it constitutes one of the major puzzles in international macroeconomics. The main goal of this chapter is to investigate the extent to which this anomaly between two crucial macroeconomic variables can be attributed to the asset component in one of them. The elimination of the nominal exchange rate with the introduction of the Euro provides a "natural" test for this inquiry. Therefore, similar as in the first part of chapter 5, the empirical analysis centers on the 12 Eurozone countries.

Chapter 5 suggested that real exchange rates appear too smooth compared to asset markets fundamentals (stochastic discount factors), while chapter 6 suggested that they appear too volatile compared to what is implied by macroeconomic fundamentals (relative consumption growth). Therefore, chapter 7 undertakes a joint analysis of both puzzles. Using the same macroeconomic dataset for the 12 Eurozone countries employed in chapter 6, the first part of the empirical analysis in chapter 7 investigates whether the (complete) elimination of the nominal exchange rate brings the level of real exchange rate volatility in line with the volatility of the underlying consumption processes. Building upon the discussion in the first part, the second part of chapter 7 further analyzes the "double disconnect" of the real exchange rate: on one hand relative to asset markets, and on the other hand relative to macroeconomic (consumption) fundamentals. For this purpose, the macroeconomic component of the real exchange rate (inflation differential) is analyzed separately from the asset component (nominal exchange rate). Constrained partly by data availability and data quality issues, the empirical analysis in this part focuses on the USA-UK country-pair over (a fairly long) time period 1975-2006.

Finally, chapter 8 gives an overview of the main findings in the empirical chapters, provides a short, over-arching conclusion, and closes by giving some suggestions for further research.

Chapter 2

Workers' Remittances and International Risk-Sharing

2.1 Introduction

One of the central benefits that the process of international financial integration offers to the residents of different countries is the possibility to diversify their macroeconomic risks internationally. Therefore, through the process of cross-border trade in assets, these countries can relax the link between domestic output growth and domestic consumption (income) growth up to the point when the latter will depend exclusively on the "world" output growth (Backus, Kehoe, and Kydland, 1992; Obstfeld and Rogoff, 1996, Chapter 5).[1] This process, through which country-specific risks are diversified away across national borders, is known as international risk-sharing. Moreover, the finance literature usually associates it with the underlying trade in financial assets. Therefore, investment in an internationally-diversified portfolio is identified as the major channel through which the process of international risk-sharing takes place. Similarly, the deviation from the hypothesis of "perfect" or complete risk-sharing is associated with the tendency by investors to "over-invest" in domestic assets and thereby forego many diversification opportunities available through investment in foreign assets. This latter phenomenon is also known

[1] Since this world output growth is the same for all countries, it implies that in the end all countries' consumption growth rates will be equalized.

in the finance literature as "home equity (or bond) bias". Since the bias in investment strategies seems to be the most obvious reason for the deviation from complete risk-sharing, many empirical studies investigate the relationship between the two.

Though the association between the two phenomena is very clear for the group of advanced economies, it might not be as important for the developing world. The reason is that portfolio investments represent a very limited portion of total capital flows for these countries. In turn, international portfolio diversification might be a relatively less important vehicle for international risk-sharing in developing compared to advanced economies. Therefore, this chapter concentrates on an *alternative channel* through which one country can smooth consumption and diversify its idiosyncratic risks internationally.

In fact, a large body of literature discusses the crucial role that changes in the current account play in (consumption) smoothing of country-specific shocks (see Sachs, 1982; Baxter, 1995; Glick and Rogoff, 1995; or Obstfeld and Rogoff, 1995, for example). In turn, the current account represents a sum of several components (balances): balance of trade, net factor incomes, and net transfer payments, implying that each of them might play a role in international consumption smoothing.[2] Furthermore, (net) transfer payments from abroad constitute an especially important portion of the current account for many developing countries (see IMF, 1993; 2004). Therefore, the focus in this empirical study is on a very specific type of transfer payments - workers' remittance flows to developing countries. In particular, the main objective is to find out what role workers' remittances play in international risk-sharing. Do countries that receive above average levels of remittance inflows achieve a significantly higher level of international risk-sharing?

The rest of this chapter is organized as follows. Section 2.2 presents a brief literature review on some important studies on international risk-sharing and workers' remittances, respectively. It is followed by section 2.3 that deals with data issues, and section 2.4 which puts forward the empirical strategy. The estimation results are presented in section 2.5. Finally, section 2.6 gives some concluding remarks and suggestions for further research.

[2]For the importance of (borrowing/lending though) foreign direct investment as a channel of international risk-sharing for financially constrained countries see Albuquerque (2003).

2.2 Literature Review

2.2.1 International Risk-Sharing

International portfolio diversification and international risk-sharing are closely interrelated, but by no means equivalent phenomena. The deviations from an internationally diversified portfolio are studied in the finance literature and are known as the "home (equity) bias", while the departures from complete consumption and income smoothing are known in the international macroeconomics literature as the "international risk-sharing puzzle". Though these concepts are interrelated, the former does not always imply the latter (nor vice-versa). More concretely, Lewis (1999) argues that "home bias" may not necessarily lead to lower international risk-sharing if most of consumption and income smoothing is done through international borrowing and lending rather than portfolio holdings. Conversely, Baxter and Jermann (1997) show that even complete international diversification of portfolio holdings does not directly imply smooth consumption and income streams. First, total equity holdings might represent a very small portion of global GDP, and second, they might not provide adequate hedging of returns to human capital.

Since the theoretical literature does not provide an unambiguous answer about the relationship between "equity home bias" and international risk-sharing, Sorensen et al. (2007) test it empirically using data for the OECD countries during the 1990s. The central issue that they investigate is whether countries with lower, or decreasing home bias during the period considered, have higher or increasing levels of international risk-sharing as well. They find that lower "home equity bias" indeed leads to more risk-sharing in consumption and income. This effect seems to be economically significant as well: a one percentage point decrease in "home bias" leads to almost half of a percentage point increase in risk-sharing.

In an influential paper, Lewis (1999) undertakes a joint investigation of the "equity home bias" as referred to in the finance literature with the "consumption home bias" as discussed in the international macroeconomics literature. Although these two phenomena seem to be influenced by similar factors, they cannot be considered equivalently. Furthermore, she invokes the fallacy in the "casual intuition" that home bias in equities and risk-sharing in consumption are necessarily linked with each other. First, equity home bias is not a *sufficient* condition for consumption home bias. In particular, as long as countries can borrow and lend with each other, consumption growth rates can be perfectly correlated even if domestic residents do not hold foreign equity at all. Second, equity home bias is not a *necessary* condition for consumption home bias. Because parts of total domestic output cannot be (are not) securitized and traded on the stock markets, perfect international portfolio diversification does not necessarily imply perfect international risk-sharing in consumption.

2.2.2 Workers' Remittances

Workers' remittances to developing countries have been steadily increasing in the past decades (see Figure 2.1). In the mid-1990s they overtook the total of private portfolio flows, thereby becoming the second most important source of foreign exchange for the developing countries (second only to FDI). Moreover, their importance as percentage of total GDP displays a similar picture (see Figure 2.2). Unambiguously, workers' remittances have become a major source of financing for many households in the developing world. Synchronously with their increasing trend, they began to capture the attention of an ever-increasing number of researchers and (international) organizations around the world.

One of the most recent contributions in this area is given in Chapter II of the World Economic Outlook prepared by the International Monetary Fund (April 2005). This analysis is threefold. First, the IMF presents some stylized facts, demonstrating that workers' remittances constitute the second largest source of foreign capital for the developing world as a whole.[3] Second, emphasis is put on the role of remittances in economic development. In this sense, they are shown to be associated with more investment in infrastructure and faster human capital accumulation. Finally, the report points out several characteristics, which make remittances especially important financing flows to developing countries. More precisely, workers' remittances are the most stable source of external finance, do not display high levels of pro-cyclicality

[3] Moreover, for many countries (especially in the Caribbean region) workers' remittances are the largest source of foreign capital and foreign exchange, representing even more than 5% of GDP for some countries.

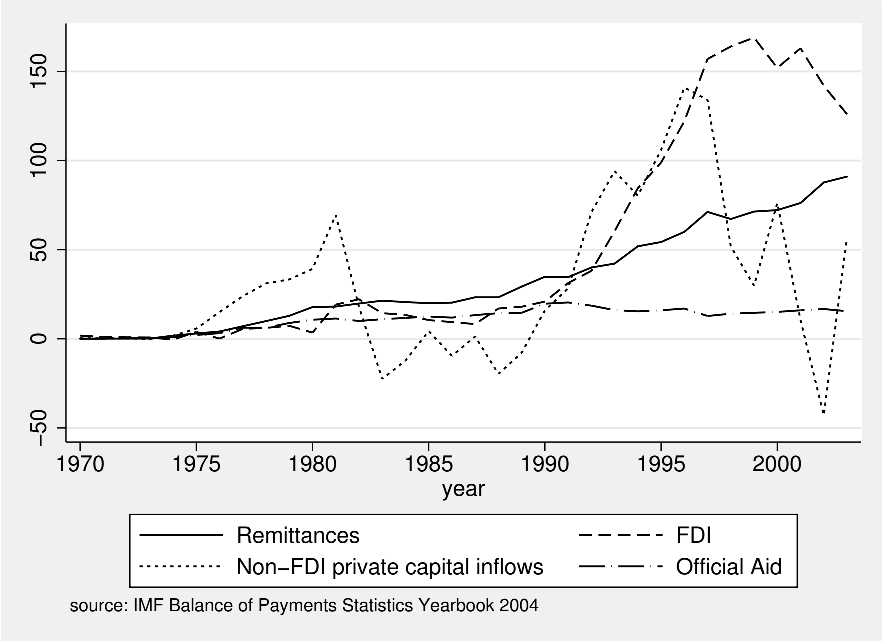


Figure 2.1: Capital Flows to Developing Countries (billions of US dollars)
Note: This figure presents capital inflows to developing countries measured in billions of US dollars. Annual observations over the period 1970-2003. All data-series come from the IMF *Balance of Payments Statistics Yearbook* 2004 and the IMF *World Economic Outlook* (April, 2005).

(see Figure 2.3), and jump sharply after an economic (or financial) crisis hits the home recipient country.[4] In fact, these last three characteristics suggest that remittances might contribute to consumption smoothing, and improve the risk-sharing capacity of the recipient economy.

The important role that remittances play in relaxing the external constraints of many developing countries is also acknowledged in Ratha (2003), published as Chapter VII of the Global Development Finance (2003) by the World Bank. Besides documenting stability as a source of external funding, Ratha (2003) distinguishes between remittances intended for consumption and remittances intended for investment.[5] Moreover, he argues that the

[4]Examples of economic crises followed by sharp increases in workers' remittance flows cited by the IMF include : Indonesia 1997, Ecuador 1999, and Argentina 2001.

[5]This distinction is mainly based on the nature of the final recipients, the flows going to households being classified as remittances intended for consumption.

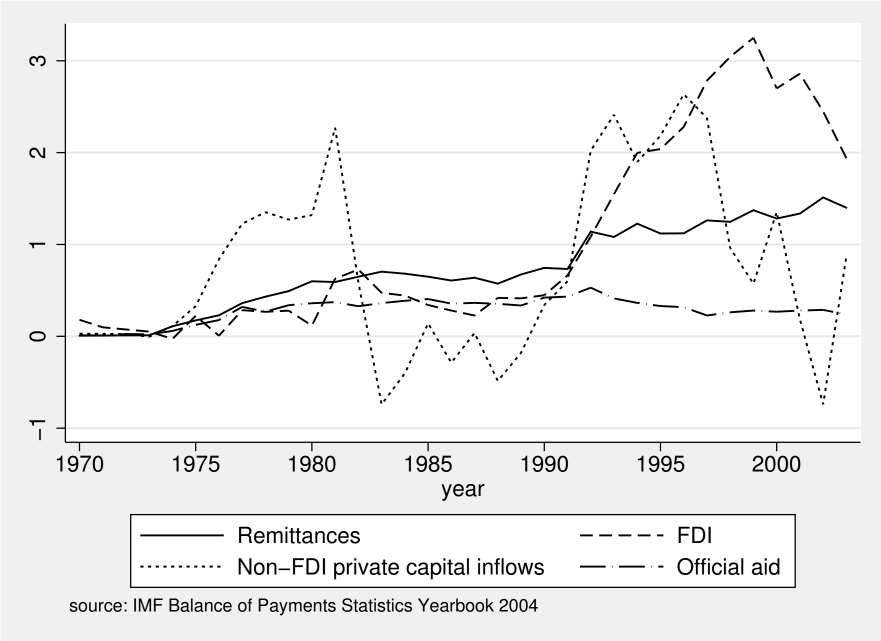


Figure 2.2: Capital Flows to Developing Countries (percent of GDP)

Note: This figure presents capital inflows to developing countries measured as percentage of GDP. Annual observations over the period 1970-2003. All data-series come from the IMF *Balance of Payments Statistics Yearbook* 2004 and the IMF *World Economic Outlook* (April, 2005).

former group should be less volatile than the latter. Additionally, he makes a distinction between different country groups. In particular, he demonstrates that poor countries with lower-than-median and middle income countries with higher-than-median growth rates receive relatively more remittances.[6]

Focusing on the macroeconomic nature of remittances, Buch and Kuckulenz (2004) find that they are mostly driven by market forces, though social and demographic considerations play an increasingly important role as well. Moreover, they find that workers' remittances are positively correlated with official capital flows (which in turn are positively correlated with private capital flows), but uncorrelated with private capital flows. Therefore, as long as official flows act as tools for macroeconomic stabilization, remittances

[6]The basic conclusion drawn from this observation is that remittances mainly serve for consumption purposes in the former and for investment purposes in the latter group.

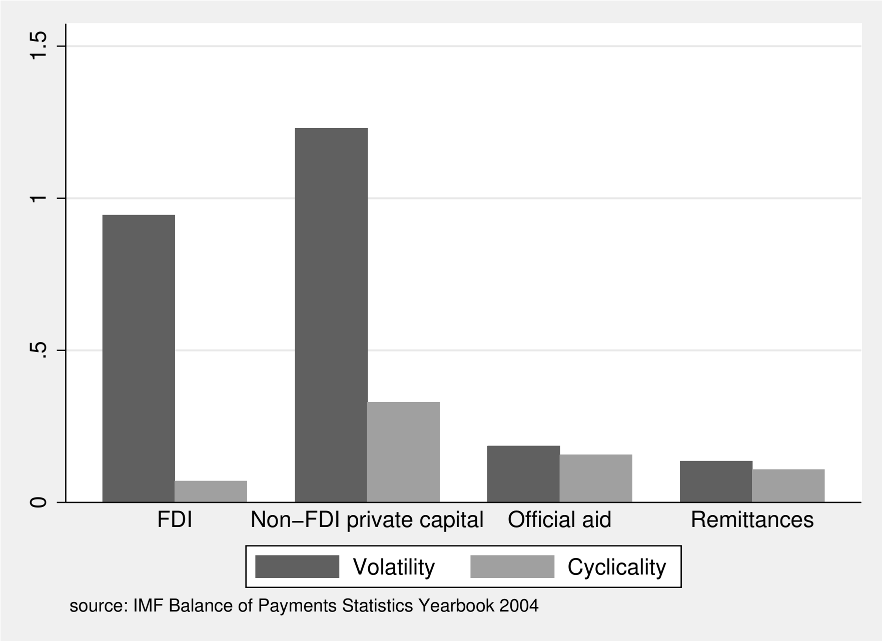


Figure 2.3: Volatility and Cyclicality of Capital Flows

Note: The figure presents volatility and cyclicality measures for the major capital inflows to developing countries over the period 1980-2003. All observations used in the calculations are annual. Volatility is defined as the standard deviation for the ratio of the corresponding capital inflow over GDP, while cyclicality is measured as the correlation between the (detrended) capital inflow and (detrended) GDP. All data-series and calculations come from the IMF *Balance of Payments Statistics Yearbook* 2004 and the IMF *World Economic Outlook* (April, 2005).

might further amplify their stabilizing or smoothing role. Concentrating on the link between financial crises and workers' remittances, Bugamelli and Paterno (2005) demonstrate that higher levels of remittances (as percentage of GDP) lead to a lower probability of a current account crisis. They attribute this to the stability of workers' remittances as a source of external financing in developing countries and their low levels of cyclicality relative to other private capital flows. Aggarwal, Demirguc-Kunt, and Martinez Peria (2006) show that workers' remittances promote financial development in developing countries by increasing the level of total deposits and credit intermediated locally. In this way, they might have a(n) further (in)direct effect on economic growth and poverty reduction. Finally, Giuliano and Ruiz-Arranz (2006) show that the effect of workers' remittances on economic growth is not uniform across countries and depends on the level of the (recipient country) financial development. In fact, the effect is the strongest for countries at low levels of financial development: workers' remittances act as an alternative financing source, and thereby, help in overcoming liquidity constrains in these countries. This section described several reasons why workers' remittances might be associated with higher international risk-sharing in consumption. The following section deals with some data issues and prepares the ground for the empirical investigation.

2.3 Data

2.3.1 Data Samples

The empirical analysis is based on an unbalanced panel dataset that includes observations on 143 countries during the period 1960-2000.[7] Due to the considerable degree of heterogeneity in terms of countries included, and the availability of data, the panel is strongly unbalanced. Thus, the number of yearly observations per country varies between 2 and 41 in the largest dataset included in the calculations, and between 2 and 11 yearly observations in the dataset that only includes developing countries in the period 1990-2000.[8]

Moreover, the total dataset is divided into several subsamples. The primary distinction is made between industrial and developing countries. The first subsample is a balanced panel that contains data on 22 industrial countries,[9] while the second subsample contains data series on 121 developing countries.[10] The latter dataset is further restricted with respect to the availability of remittances data: the total number of developing countries decreases to 117 in specifications that include the remittances term. Finally, various subsam-

[7]Data-series were available for 153 countries in total. However, due to the large number of missing observations for 10 countries, the final dataset is restricted to 143 countries.

[8]Description of the complete dataset including list of countries and data-series specification is available upon request.

[9]This dataset represents a balanced panel as it contains 875 yearly observations for 22 countries over the period 1960-2000.

[10]More precisely, this subsample refers to a rather heterogeneous group of countries that includes all countries not included in the subsample of industrial countries. This dataset is strongly unbalanced.

ples of developing countries are used in some of the estimations. The entire sample of 117 developing countries is divided into the following subsamples: more financially integrated countries (MFIs) (20 countries), less financially integrated countries (83 countries), and transition economies (14 countries).[11]

2.3.2 Data Sources

There are two types of data used in the construction of the panel dataset: macroeconomic data and data on workers' remittances. In turn, they come from two different sources: the macroeconomic series are taken from the Penn World Table Version 6.1 and the data on remittances comes from the IMF Balance of Payments Statistics Yearbook.[12] The first group refers to the following series: real consumption per capita and real GDP per capita, both adjusted for PPP, which are standardly used in macroeconomic studies. For the purposes of this study, we calculate the yearly growth rates of these two series for each country available and for the world as a whole. The growth rates for global GDP and global consumption are calculated as the unweighed average of the real GDP per capita growth rates for the set of industrial countries in the corresponding year.[13] We include only the group of industrial countries in the calculation of the "global" figures because this group constitutes the largest portion of both world GDP and world consumption.[14]

[11]The classification of developing countries into the first two groups MFIs and LFIs is done by the IMF. The group of transition economies includes the former centrally-planned economies with the exception of the countries that belong to the Commonwealth of Independent States (ex-Russia) and Bosnia-Herzegovina and FR Yugoslavia (later Serbia and Montenegro). Therefore, this groups includes the following 14 countries: Albania, Bulgaria, Croatia, Czech Republic, Estonia, Hungary, Latvia, Lithuania, Macedonia, Poland, Romania, Russia, Slovakia, and Slovenia.

[12]Special thanks to Nicola Spatafora and Angela Espiritu from the IMF for providing the dataset on workers' remittances.

[13]The estimation results for different subsamples do not change significantly when "world" growth rates are calculated as unweighed averages of the growth rates for that particular subsample only.

[14]In fact, including the diverse group of developing countries in the calculation of global figures might include additional ambiguities. For example, there are large differences in the economic weights attached to each country, depending on whether they are based on market-exchange rates (as usually calculated by the World Bank), or on purchasing power parity rates (as calculated by the IMF and the OECD).

2.3.3 Remittances Data

In general, data on remittances is inaccurate and unreliable. Moreover, it is not always comparable across countries because different sources (national agencies, central banks, statistical offices, international bodies, etc.) use different methodologies and variable definitions in their compilations.[15] In this respect, the Balance of Payments Statistics Yearbook by the IMF, used in this study, is generally acknowledged as the most complete and reliable database on workers' remittances available.[16] Since there is not an unambiguous or universal definition of workers' remittances, various variables are used in the empirical literature. Generally, the Balance of Payments Statistics Yearbook reports three components that can be used for this purpose: compensation of employees, workers remittances, and migrants' transfers.[17] The first two categories are registered with the current account, while the last category belongs to the capital account of the balance of payments.[18] More precisely, *compensation of employees* refers to "the wages, salaries, and other benefits earned by individuals for work that they performed in economies in which they are not residents", while *workers' remittances* "cover current transfers by migrants who work and are considered residents of new economies (other than the transfers' final destination)". Therefore, the slight distinction between these two components is based on the time period that the migrant (remitter) is expected to stay in the new economy.[19] Finally, *migrants' transfers* refers to the flows of (financial) assets that are related to persons who migrate from one economy to another.

In this study remittances are defined as the sum of all three components.

[15] For more on the difficulties in cross-country comparisons for remittances data and the efforts made to improve its quality see de Luna Martinez (2005).

[16] See Reinke and Patterson (2005) and the IMF Balance of Payments Manual (5th ed.) for more details on definitions and data-related issues.

[17] For some countries, additional entries registered under "current private transfers" refer specifically to remittances.

[18] The definitions given here follow Reinke and Patterson (2005).

[19] By convention, migrants that stayed or are expected to stay in the new economy for more than one year are considered residents, and therefore, their transfers are included in the latter category. If the workers are expected to stay less then one year, their transfers are treated as compensation of employees.

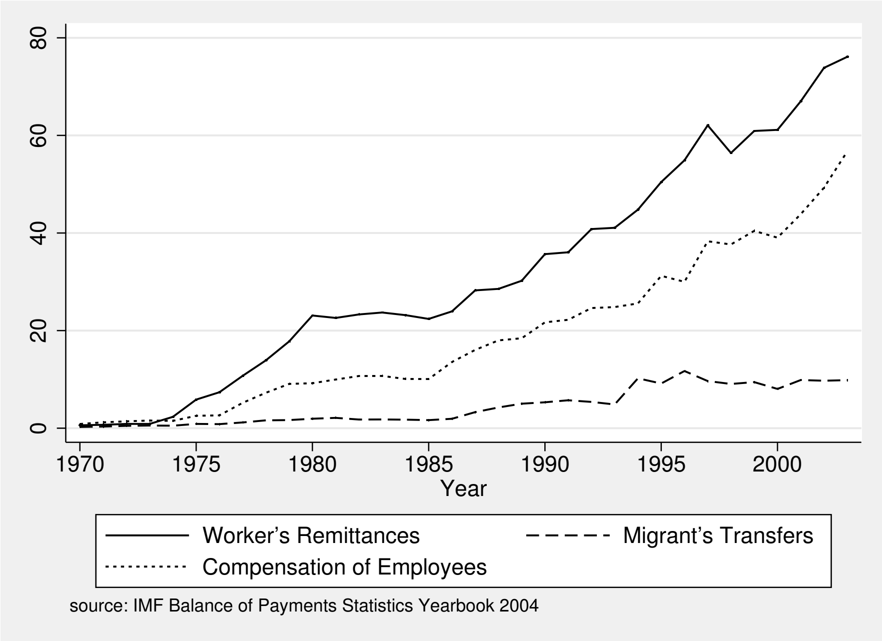


Figure 2.4: Remittances Components (billions of US dollars)

Note: The figure presents the evolution in the components of workers' remittances flows to the developing countries during the period 1970-2003. All observations are annual and measured in billions of US dollars. All data-series come from the IMF *Balance of Payments Statistics Yearbook* 2004.

This choice can be justified on several grounds. First, economic theory does not give clear guidelines about the appropriate definition. Second, the distinctions among the three components are small (and one may argue arbitrary), and therefore the probability for misclassification of certain transfers is non-negligible. Finally, though large differences exist among the three components, Figure 2.4 shows that there was an increasing trend in all of them during the past four decades.

An important point to note is that only gross data on remittances inflows have been used in the construction of workers' remittances indicators. There are at least three reasons why gross and not net (inflows minus outflows) might be more appropriate for this analysis. First, workers' remittances constitute an especially important source of external financing for the developing world, where countries are most often only recipients of these funds. Second, the data on remittance outflows is even more scarce, of worse quality, and less reliable than data on remittance inflows. Third, most countries report data either on inflows or on outflows, but not on both of them. Therefore, it would be very difficult to calculate precise and meaningful indicators for net remittance flows. One of the major disadvantages of this strategy is that the empirical analysis of remittance flows is limited to the set of developing countries.

2.4 Empirical Specification

There are two main issues that we investigate in this study. First, we present some measures of the degree of international risk-sharing in consumption for different groups of countries during the period 1960-2000. Second, we attempt to shed light on the major question of this chapter: what is the role of workers' remittances in international risk-sharing?

2.4.1 Simple Risk-Sharing Tests

In order to measure the degree of risk-sharing in consumption we estimate the following type of panel regression equations:[20]

$$(\Delta \log C_{it} - \Delta \log C_t) = \alpha + \beta(\Delta \log GDP_{it} - \Delta \log GDP_t) + \epsilon_{it} \qquad (2.1)$$

In this equation $\Delta \log C_{it}$ is the year-on-year growth rate of real consumption per capita for country i in year t, $\Delta \log C_t$ is the growth rate for "world" real consumption per capita, $\Delta \log GDP_{it}$ and $\Delta \log GDP_t$ are the corresponding terms for GDP, β measures the average co-movement of the countries' idiosyncratic consumption growth with their idiosyncratic GDP growth during the entire time period, and ϵ is the error term.

The slope coefficient β deserves special attention because it measures the (average) deviation from perfect risk-sharing in consumption. In particular, the perfect international risk-sharing in consumption hypothesis states that if the countries manage to share completely the idiosyncratic risks that they face, then this coefficient should not be significantly different from zero.[21] In a

[20] The empirical specifications used in this study closely follow Sorensen et al. (2007).

[21] Strictly speaking, the null hypothesis of perfect international risk-sharing also implies that the constant term α should not be significantly different from zero.

corresponding manner, one may argue that $100(1-\beta)\%$ measures the degree of international risk-sharing in percentage terms.

2.4.2 Augmented Specification

A similar type of regressions is estimated to measure the effect of workers' remittances in international risk-sharing. In particular, we estimate the following panel regressions:

$$(\Delta \log C_{it} - \Delta \log C_t) = \alpha + \psi(\Delta \log GDP_{it} - \Delta \log GDP_t) + \epsilon_{it} \tag{2.2}$$

where the slope coefficient is defined as follows:

$$\psi = \psi_0 + \psi_1(t-\bar{t}) + \psi_2 \log\left(\frac{R_{it}}{\overline{R}_t}\right) \tag{2.3}$$

$\bar{t}$ is the middle year of the sample period (i.e. 1995 when the sample refers to the period 1990-2000), R_{it} is the ratio between total remittances received and GDP for country i in year t and $\overline{R}_t$ is the average ratio between total remittances received and GDP in year t. Therefore, $t-\bar{t}$ can be thought of as a time trend, which captures the trend decline in consumption smoothing not directly caused by the increase in workers' remittances. Finally, the term $\log\left(\frac{R_{it}}{\overline{R}_t}\right)$ measures the relative importance of the ratio total remittances received to GDP for a certain country i in year t compared to its average value across countries in year t.

Defined in this way, each of the coefficients ψ_0, ψ_1 and ψ_2 has very precise meaning. In particular, $1-\psi_0$ measures the degree of international consumption risk-sharing achieved by country with the "average" ratio between remittances received and GDP during the middle year $\bar{t}$. By similar argument, the coefficient $-\psi_1$ gives the average year-on-year increase in consumption risk-sharing. If it is true that country-specific risks became better diversified internationally through time, then one will expect a negative sign for ψ_1. Furthermore, $-\psi_2$ captures the impact of a higher than average ratio between total remittances received and GDP for a certain country on its ability to smooth idiosyncratic output shocks. Therefore, a significantly negative value for ψ_2 implies better risk-sharing for countries that receive above-average ratio of workers' remittances relative to GDP. Finally, the entire coefficient $1-\psi = 1-\psi_0-\psi_1(t-\bar{t})-\psi_2 \log\left(\frac{R_{it}}{\overline{R}_t}\right)$ measures the amount of consumption risk-sharing achieved by country i in year t.

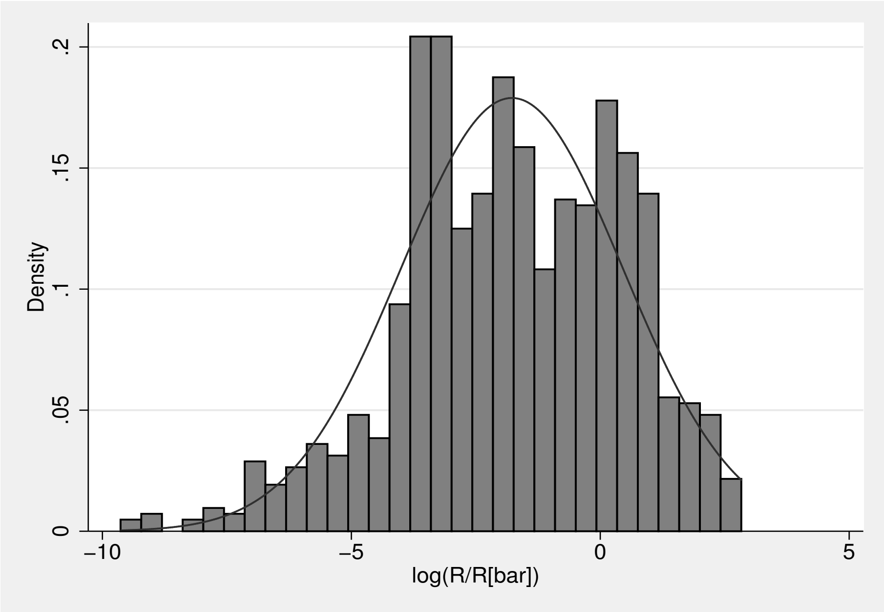
2.4.3 Remittances Indicator

It is important to note that alternative definitions for the remittances term $\log\left(\frac{R_{it}}{\overline{R}_t}\right)$ are employed in the empirical analysis. In fact, the denominator in this term - the average value for the ratio of workers' remittances over GDP ($\overline{R}_t$) can be calculated in alternative ways. For example, it can be calculated as the ratio between the average level of remittances at time t and the average GDP level at time t: $\overline{R}_t = \overline{Rem}_t/\overline{GDP}_t$. Alternatively, it can be calculated as the typical or average ratio across all countries at time t: $\overline{R}_t = \overline{(Rem_t/GDP_t)}$. In general, $\overline{R}_t$ calculated as the ratio between averages (first case) will differ from $\overline{R}_t$ calculated as average ratio (second case). In turn, the differences in the denominator will be translated into differences in the remittances indicator included in the empirical specification.

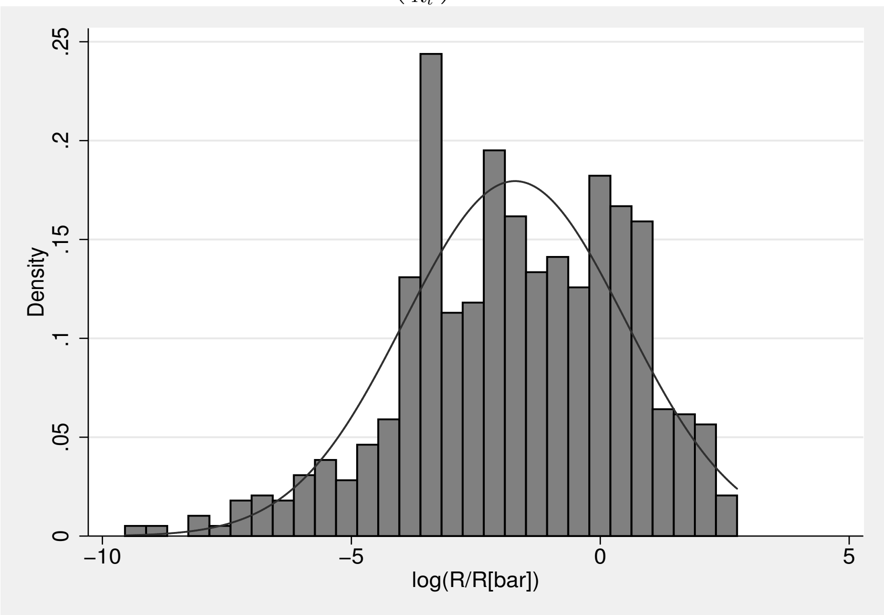
Figure 2.5 presents two histograms for the remittances indicator $\log\left(\frac{R_{it}}{\overline{R}_t}\right)$, based on alternative definitions for $\overline{R}_t$.[22] In the upper graph $\overline{R}_t = \overline{Rem}_t/\overline{GDP}_t$, while in the lower $\overline{R}_t = \overline{(Rem_t/GDP_t)}$. Although the two indicators seem to have very similar distributions, they are not identical. To control for the possibility that the estimation results depend on the particular definition of the remittances term, all empirical specifications are estimated using both remittances indicators.[23]

[22]In both cases, the indicator $(R_{it}/\overline{R}_t)$ displays a very skewed distribution with an average value of 1. Therefore, taking the natural logarithm improves normality. Both histograms in Figure 2.5 refer to the natural logarithm of this indicator. The results do not crucially depend on whether $(R_{it}/\overline{R}_t)$ or its logarithm is included in the regressions.

[23]The specifications that use the first definition for the remittances indicator are marked as "Remittances(m)", while those that use the second definition are marked as "Remittances(t)".



a) Distribution of $\log\left(\frac{R_{it}}{\overline{R}_t}\right)$ when $\overline{R}_t = \overline{Rem}_t/\overline{GDP}_t$



b) Distribution of $\log\left(\frac{R_{it}}{\overline{R}_t}\right)$ when $\overline{R}_t = \overline{(Rem_t/GDP_t)}$

Figure 2.5: Alternative Remittances Indicators (Histogram)

2.5 Results

2.5.1 Simple Risk-Sharing Tests

The estimation results for the panel specification in equation (2.1) are given in Table 2.1. Similar type of equations is estimated for the subsamples of developing and industrial countries during the period 1960-2000. Moreover, each equation has been estimated with fixed effects (assuming non-zero correlation between the regressors and the country-specific part of the error term) and with random effects panel data estimation techniques.[24] As can be seen in Table 2.1, the results obtained using the two alternative estimation techniques do not differ a lot. Moreover, the Hausman specification test suggests that the difference between coefficients is not significantly different from zero. Thereby, this specification test implies that both estimation techniques (fixed-effects and random-effects) yield consistent estimates, while the random-effects estimates are also efficient in this case.[25]

Table 2.1 displays several important findings. First, in line with many previous studies, the hypothesis of perfect risk-sharing in consumption is rejected for each (sub)sample of countries. Moreover, the results imply large differences across the two groups of countries. In particular, the group of developing countries achieves a "relatively" low degree of international risk-sharing compared to the group of advanced economies. In fact, about 17.3% $(1 - 0.827)$ of country-specific risks for the group of developing countries is shared internationally.[26] Conversely, idiosyncratic consumption changes are much less dependent on idiosyncratic GDP changes in the group of advanced economies. If one performs similar calculations for the group of industrial countries, the percentage of risks shared internationally is 40% $[100(1 - 0.60)\%]$.

[24]The adjustment of the standard errors for intragroup (intertemporal) serial correlation produces somewhat lower t-values, but does not lead to significant changes in the main results.

[25]In most specifications in this empirical study, the null hypothesis that there is no systematic difference between the coefficients obtained with fixed-effects and the coefficients obtained with random-effects cannot be rejected. Therefore, results from both estimation procedures are reported in the tables.

[26]The degree of international risk-sharing in percentage terms is given by $100(1 - \beta)\%$, where β is the slope coefficient estimate.

Table 2.1: International Risk-Sharing 1960-2000

	Developing Countries		Industrial Countries	
	RE	FE	RE	FE
Constant	0.0013	0.0013	0.00067	0.00068
	(0.98)	(0.97)	(0.55)	(0.55)
Output	0.827	0.826	0.604	0.587
	(63.16)***	(61.30)***	(14.87)***	(13.93)***
Observations	3898	3898	875	875
Countries	121	121	22	22
R-squared		0.51		0.21

Note: The table presents results from panel data estimations based on yearly observations over time period 1960-2000 for developing and industrial countries. The empirical specification is given by the following regression equation: $(\Delta \log C_{it} - \Delta \log C_t) = \alpha + \beta(\Delta \log GDP_{it} - \Delta \log GDP_t) + \epsilon_{it}$. Dependent variable is $(\Delta \log C_{it} - \Delta \log C_t)$ - the idiosyncratic real per capita consumption growth rate for country i at time t. The columns named RE and FE in the table refer to random-effects and fixed-effects estimation results, respectively. For each specification, the table displays the slope coefficient estimates with the corresponding t-statistics (in brackets). Significance at 10%, 5%, and 1% is indicated by *, **, and ***, respectively. The critical values for the t-distribution are 1.645 (10%), 1.960 (5%), and 2.576 (1%).

2.5.2 Risk-Sharing in Developing Countries

Tables 2.2, 2.3 and 2.4 display estimation results for the non-linear specification given by equations (2.2) and (2.3). This non-linear specification is of central importance in this study as it shows the effect of workers' remittances on international consumption smoothing. The analysis is limited to the time-period 1990-2000 because the data on remittances for the period 1960-1990 is very scarce and available for a small number of countries only.

Table 2.2 presents results for the entire sample of developing countries during the period 1990-2000. Three findings are worth mentioning. First, the slope coefficient of the idiosyncratic output growth term declines in each specification, indicating that the "average" international risk-sharing among developing countries increases compared to the baseline specification (Table 2.1).[27] Depending on the specification employed in Table 2.2, the degree of in-

[27] The "average" refers to the country with an average ratio remittances received to GDP during the "middle" year, i.e. 1995 in this case.

Table 2.2: International Risk-Sharing in Developing Countries 1990-2000

	RE	FE	RE	FE	RE	FE
Constant	-0.001	-0.001	0.003	0.002	0.003	0.002
	(-0.23)	(-0.34)	(1.00)	(0.75)	(1.01)	(0.77)
Output	0.773	0.754	0.753	0.742	0.75	0.738
	(29.27)***	(26.79)***	(21.89)***	(20.52)***	(21.20)***	(19.88)***
Time-trend			-0.032	-0.025	-0.032	-0.025
			(-3.13)***	(-2.29)**	(-3.09)***	(-2.25)**
Remittances(m)					-0.025	-0.025
					(-2.11)**	(-1.99)**
Remittances(t)			-0.024	-0.024		
			(-2.10)**	(-1.97)**		
Observations	1218	1218	997	997	997	997
Countries	121	121	117	117	117	117
R-squared		0.4		0.42		0.42

Note: The table presents results from panel data estimations based on yearly observations over time period 1990-2000 for the entire group of developing countries. The results refer to three alternative specifications. The first two columns refer to a simple risk-sharing test without interaction terms given by the following regression equation: $(\Delta \log C_{it} - \Delta \log C_t) = \alpha + \beta(\Delta \log GDP_{it} - \Delta \log GDP_t) + \epsilon_{it}$. The other four columns present results from non-linear specifications that include time-trend and remittances interaction terms: $(\Delta \log C_{it} - \Delta \log C_t) = \alpha + \psi(\Delta \log GDP_{it} - \Delta \log GDP_t) + \epsilon_{it}$, where $\psi = \psi_0 + \psi_1(t - \bar{t}) + \psi_2 \log(R_t/\overline{R_t})$. The last two specification differ with respect to the definition of the remittances indicator used: $\overline{R_t} = \overline{Rem_t}/\overline{GDP_t}$ in the first, and $\overline{R_t} = \overline{(Rem_t/GDP_t)}$ in the second specification. Dependent variable in all specifications is $(\Delta \log C_{it} - \Delta \log C_t)$ - the idiosyncratic real per capita consumption growth rate for country i at time t. The columns named RE and FE in the table refer to random-effects and fixed-effects estimation results, respectively. For each specification, the table displays the slope coefficient estimates with the corresponding t-statistics (in brackets). Significance at 10%, 5%, and 1% is indicated by *, **, and ***, respectively. The critical values for the t-distribution are 1.645 (10%), 1.960 (5%), and 2.576 (1%).

ternational consumption risk-sharing lies in the range 22.7-26.2%.[28] However, the slope coefficient in front of the idiosyncratic output growth term is still significantly different from zero at any conventional significance level, indicating strong violation of the hypothesis of perfect risk-sharing. Moreover, the time trend has a significantly negative slope coefficient, meaning that risk-sharing improved gradually through time. The final and most important finding refers to the "remittances interaction term" which has negative value significant at the 5% significance level.[29] This suggests that countries that receive more than average amount of workers' remittances per year (relative to their GDP) achieve a significantly higher degree of international risk-sharing in consump-

[28] These figures for "average" risk-sharing are calculated according to $100(1-\beta)\%$, where β is the slope coefficient in front of the idiosyncratic output growth term. The lowest degree is found in the baseline specification (first column) 22.7%=$100(1-0.773)\%$, while the highest is found for the last specification 26.2%=$100(1-0.738)\%$.

[29] The results are not driven by non-stationarity of the variables. In fact, the Im, Pesaran, and Shin and Madala and Wu tests both reject the null hypothesis of unit-root at 1% significance level.

Table 2.3: Risk-Sharing in MFIs, LFIs, and Transition Economies 1990-2000

	MFI		LFI		Transition Economies	
	RE	FE	RE	FE	RE	FE
Constant	-0.001	-0.001	0.004	0.002	0.012	0.011
	(-0.29)	(-0.47)	(0.80)	(0.52)	(1.26)	(1.12)
Output	0.85	0.856	0.78	0.767	0.286	0.201
	(21.57)***	(21.48)***	(17.50)***	(16.39)***	(2.17)**	(1.37)
Time-trend	0.015	0.013	-0.046	-0.038	0.076	0.100
	(1.08)	(0.91)	(-3.84)**	(-2.98)**	(1.49)	(1.78)*
Remittances(m)	-0.007	-0.003	-0.015	-0.016	-0.131	-0.150
	(-0.36)	(-0.16)	(-1.09)	(-1.05)	(-2.99)***	(-3.20)***
Observations	185	185	710	710	102	102
Countries	20	20	83	83	14	14
R-squared		0.75		0.43		0.32

Note: The table presents results from panel data estimations based on yearly observations over time period 1990-2000 for the several subgroups of developing countries. The first two columns contain results for more financially integrated countries (MFIs), the middle two columns for the group of less financially integrated countries (LFIs), and the last two columns for the groups of 14 transition economies. The (non-linear) empirical specification is given by the following regression equation: $(\Delta \log C_{it} - \Delta \log C_t) = \alpha + \psi(\Delta \log GDP_{it} - \Delta \log GDP_t) + \epsilon_{it}$, where $\psi = \psi_0 + \psi_1(t - \bar{t}) + \psi_2 \log(R_t/\overline{R}_t)$. The remittances indicator is defined as the average level of remittances received at time t over the average level of GDP at time t: $\overline{R}_t = \overline{Rem}_t/\overline{GDP}_t$. Dependent variable in all specifications is $(\Delta \log C_{it} - \Delta \log C_t)$ - the idiosyncratic real per capita consumption growth rate for country i at time t. The columns named RE and FE in the table refer to random-effects and fixed-effects estimation results, respectively. For each specification, the table displays the slope coefficient estimates with the corresponding t-statistics (in brackets). Significance at 10%, 5%, and 1% is indicated by *, **, and ***, respectively. The critical values for the t-distribution are 1.645 (10%), 1.960 (5%), and 2.576 (1%).

tion as well.[30]

The analysis goes one step further in Table 2.3 and Table 2.4. The entire set of developing countries is divided into three subgroups: more financially integrated (MFIs), less financially integrated (LFIs), and transition economies. The results for the first two groups are broadly similar to those for the entire set of developing countries. However, there are three changes that deserve particular attention. First, the time trend is not significant anymore for the MFIs and the transition economies, though it stays significantly negative for the LFIs.[31] Second, although its sign stays negative, the remittances interaction term is not significantly different from zero for the MFIs and the LFIs. On the contrary, the impact of remittances on international risk-sharing seems

[30] The two middle columns differ from the last two columns in Table 2.2 with respect to the definition of the remittances interaction term. The former employs the average ratio of workers' remittances over GDP, while the latter uses the ratio between average level of workers' remittances and average GDP level.

[31] The positive (and insignificant) time-trend for the group of MFIs might reflect the macroeconomic turbulence they faced in the aftermath of the financial crises in the 1990s.

Table 2.4: Risk-Sharing in MFIs, LFIs, and Transition Economies 1990-2000

	MFI		LFI		Transition Economies	
	RE	FE	RE	FE	RE	FE
Constant	-0.001	-0.001	0.004	0.002	0.011	0.010
	(-0.30)	(-0.47)	(0.80)	(0.51)	(1.22)	(1.08)
Output	0.851	0.857	0.782	0.77	0.306	0.224
	(21.53)***	(21.36)***	(18.07)***	(16.92)***	(2.39)**	(1.56)
Time-trend	0.015	0.013	-0.046	-0.038	0.074	0.097
	(1.07)	(0.91)	(-3.86)***	(-3.00)***	(1.44)	(1.73)*
Remittances(t)	-0.007	-0.004	-0.015	-0.016	-0.13	-0.15
	(-0.38)	(-0.18)	(-1.07)	(-1.04)	(-2.98)***	(-3.19)***
Observations	185	185	710	710	102	102
Countries	20	20	83	83	14	14
R-squared		0.75		0.43		0.32

Note: The table presents results from panel data estimations based on yearly observations over time period 1990-2000 for the several subgroups of developing countries. The first two columns contain results for more financially integrated countries (MFIs), the middle two columns for the group of less financially integrated countries (LFIs), and the last two columns for the groups of 14 transition economies. The (non-linear) empirical specification is given by the following regression equation: $(\Delta \log C_{it} - \Delta \log C_t) = \alpha + \psi(\Delta \log GDP_{it} - \Delta \log GDP_t) + \epsilon_{it}$, where $\psi = \psi_0 + \psi_1(t - \bar{t}) + \psi_2 \log(R_t/\overline{R}_t)$. The remittances indicator is defined as the average value for the ratio of remittances received over GDP at time t: $\overline{R}_t = \overline{(Rem_t/GDP_t)}$. Dependent variable in all specifications is $(\Delta \log C_{it} - \Delta \log C_t)$ - the idiosyncratic real per capita consumption growth rate for country i at time t. The columns named RE and FE in the table refer to random-effects and fixed-effects estimation results, respectively. For each specification, the table displays the slope coefficient estimates with the corresponding t-statistics (in brackets). Significance at 10%, 5%, and 1% is indicated by *, **, and ***, respectively. The critical values for the t-distribution are 1.645 (10%), 1.960 (5%), and 2.576 (1%).

to strengthen for the group of transition economies as its slope coefficient becomes even more significantly negative.[32] Finally, the slope coefficient ψ_0 in front of the idiosyncratic output growth rate strongly decreases for the group of transition economies: from about 0.70-0.80 in the results for all developing countries (Table 2.2) to only about 0.2-0.3 in this subgroup. Moreover, it turns insignificant in the fixed-effects specification for this group. This means that the null hypothesis of full international risk-sharing for the group of transition economies, once the time trend and the effect of workers' remittances is accounted for, cannot be rejected at conventional significance levels (even at 10% significance).

Finally, Table 2.4 reports estimation results from a specification similar to the one in Table 2.3, the sole difference being that the alternative definition for the remittances indicator is used.[33] As can be seen from the Table, the results

[32]The null hypothesis that this slope coefficient is not significantly different from zero can be rejected even at 1% significance level now. The corresponding significance level for the entire set of developing countries was 5%.

[33]This checks for the *sensitivity* of the main results w.r.t. the definition of this indicator.

stay virtually the same, meaning that the major conclusions with respect to the impact of workers' remittances are not sensitive to the specific definition of the remittances indicator.

2.5.3 Exclusion of Transition Economies

There are two troubling and counterintuitive findings in Tables 2.3 and 2.4 that might throw some doubts on the reliability of the estimation results for the transition economies. First, the time-trend interaction term enters the regression equation with a positive sign meaning that international risk-sharing for this group of countries decreased over time (1990-2000), despite their significant and continual economic integration over the corresponding period. Second, the coefficient in front of the idiosyncratic output growth term drops dramatically for the subgroup of transition economies, thereby implying implausibly high levels of risk-sharing.

Most data-series for the transition economies are short (starting only in 1993-1994) and of generally low quality. Looking more closely at the cross-sections per year, the results for this group appear largely driven by the starting years 1993-1995 when both consumption and output growth were characterized by erratic movements and significant measurement errors. In fact, idiosyncratic consumption and idiosyncratic output growth are especially high (in absolute terms) for most transition economies in these starting years (easily above 20 percent per year). Therefore, if included together with "normal" observations, such extreme observations might drive most estimation results. As long as they are simply products of measurement errors at the beginning of the transition, their inclusion in the analysis seems inappropriate.

In order to investigate whether the results for the group of developing countries are driven primarily by the group of transition economies, this section excludes this group from the estimations. Table 2.5 reports results from panel data estimations for the MFIs and LFIs only. There are three important findings in this table. First, the idiosyncratic output growth term enters with a significantly positive sign, confirming the evidence of strong departure from the null hypothesis of perfect international risk-sharing. Second, the time-trend enters with a significantly negative sign, reflecting the general

Table 2.5: International Risk-Sharing in MFIs and LFIs 1990-2000

	RE	FE	RE	FE	RE	FE
Constant	0.003	0.002	0.003	0.003	0.002	0.002
	(0.82)	(0.91)	(0.88)	(1.06)	(0.90)	(0.97)
Output	0.676	0.669	0.675	0.671	0.686	0.664
	(12.53)***	(11.73)***	(12.38)***	(11.69)***	(20.78)***	(18.70)***
Time-trend	-0.049	-0.041			-0.029	-0.028
	(-4.35)***	(-3.38)***			(-2.69)***	(-2.43)**
Remittances(m)	-0.017	-0.014	-0.016	-0.013		
	(-1.17)	(-0.87)	(-1.11)	(-0.84)		
Observations	895	895	895	895	1089	1089
Countries	103	103	103	103	107	107
R-squared		0.30		0.27		0.27

Note: The table presents results from panel data estimations based on yearly observations for a sample that includes all MFIs and LFIs over time period 1990-2000.The general (non-linear) empirical specification is given by the following regression equation: $(\Delta \log C_{it} - \Delta \log C_t) = \alpha + \psi(\Delta \log GDP_{it} - \Delta \log GDP_t) + \epsilon_{it}$, where $\psi = \psi_0 + \psi_1(t - \bar{t}) + \psi_2 \log(R_t/\overline{R}_t)$. The first two columns contain results from the complete specification, the middle two columns exclude the time-trend, while the last two columns exclude the remittances interaction term. The remittances indicator is defined as the average level of remittances received at time t over the average level of GDP at time t: $\overline{R}_t = \overline{Rem}_t/\overline{GDP}_t$. Dependent variable in all specifications is $(\Delta \log C_{it} - \Delta \log C_t)$ - the idiosyncratic real per capita consumption growth rate for country i at time t. The columns named RE and FE in the table refer to random-effects and fixed-effects estimation results, respectively. For each specification, the table displays the slope coefficient estimates with the corresponding t-statistics (in brackets). Significance at 10%, 5%, and 1% is indicated by *, **, and ***, respectively. The critical values for the t-distribution are 1.645 (10%), 1.960 (5%), and 2.576 (1%).

trend of increasing risk-sharing over time for these countries. Finally, the slope coefficient in front of the remittances term is consistently negative in each specification, though not significant statistically. In sum, this finding implies that higher than average levels of workers' remittances per capita are associated with relatively higher international risk-sharing, though this effect is not statistically significant once transition economies are excluded from the analysis.

The results for the empirical specifications that include the alternative definition for the remittances indicator are very similar to the ones presented in Table 2.5, and therefore, are not presented here.[34]

2.6 Conclusion

During the past four decades, idiosyncratic consumption growth rates have been very strongly correlated with idiosyncratic output growth rates, thereby

[34]The remittances term enters all specifications with a negative sign. Nonetheless, this effect is statistically insignificant with typical t-values in the range $(-0.90, -1.25)$.

suggesting that a very small part of macroeconomic risks is actually shared internationally. In this sense, the results presented in this chapter are in line with most other empirical studies on international risk-sharing that use macroeconomic data. Various explanations have been proposed for this apparent puzzle in international macroeconomics. In this respect, the reduction in equity home bias has been identified as a crucial channel through which risk-sharing across countries can be improved. This chapter offered an alternative risk-sharing channel - workers' remittances sent to their home countries. Moreover, the empirical evidence gave support to this alternative channel. Indeed, developing countries with above average remittance receipts during the last decade of the previous century are associated with smaller deviations from the perfect risk-sharing hypothesis. Though this relationship seems to be true for each subgroup of developing countries (MFIs, LFIs, and transition economies), it is statistically significant only for the transition economies. In fact, the amplification of this effect for the transition economies might be just a manifestation of the poor data quality for this group of countries. This is especially true in view of the (spurious) time-trend coefficient found for the transition countries, indicating decreasing risk-sharing through time. Therefore, we conclude that higher than average levels of workers' remittances per capita do not significantly raise international risk-sharing for the group of developing countries.

Apart from presenting empirical evidence on an alternative risk-sharing channel, this study suggests several questions that deserve further research. First, it is important to find out why the impact of workers' remittances might be different across different groups of developing countries. In turn, identifying the underlying reasons might give further insight into the broad institutional environment through which workers' remittances arrive to their final beneficiaries. Finally, the increasing trend in remittance inflows might be accompanied by additional private (and/or official) capital flows to developing countries. Thereby, remittance inflows might fundamentally change the entire framework through which these countries diversify their macroeconomic risks.

Chapter 3

The Importance of Estimation Bias in Conventional Tests of Uncovered Interest Parity

3.1 Introduction

Uncovered interest rate parity (UIP) is an equilibrium condition that binds together the (expected) returns on two comparable assets denominated in different currencies. It suggests that higher interest rate currencies should depreciate ex-post. Thereby, it provides a crucial theoretical underpinning for many models in international finance and international monetary economics.

Nonetheless, the empirical evidence is generally unfavorable to UIP. In fact, most empirical tests soundly reject this condition, and typically find significant coefficients with the wrong sign. Hence, they suggest that higher interest rate currencies tend to appreciate (not depreciate) ex-post. The literature has offered a number of reasons that (partially) account for this empirical "anomaly", ranging from time-varying nature of risk-premia to sophisticated econometric properties in small samples or extreme observations (Fama, 1984; Flood and Rose, 1996; Huisman, Koedijk, Kool, and Nissen, 1998; Baillie and Bollerslev, 2000; Bekaert, Wei, and Xing, 2006).

In this study we investigate an alternative explanation. Following the approach employed in Kool and Thornton (2004) and Thornton (2006), we show that the empirical procedure conventionally used to test for UIP may produce biased slope coefficients if the true data-generating process slightly differs from the theoretically expected one. Moreover, this bias crucially depends on the correlation between the two interest rate series and on the relative interest rate volatility of the countries used in the estimations. It obtains the largest values (in absolute terms) when interest rate series are strongly correlated and display similar volatility levels. The central aim of this chapter is to empirically investigate the importance of this estimation bias on slope coefficient estimates obtained in conventional tests of UIP. We do not attempt to "rehabilitate" the UIP nor solve the UIP puzzle, but rather provide evidence about one problem (estimation bias) related to the conventional UIP testing procedure.

The rest of this chapter is organized as follows. Section 3.2 presents the concept and the conventional empirical procedure used to test for UIP, and formally demonstrates why slope coefficient estimates from these conventional tests of UIP may be biased. Section 3.3 deals with data issues and the empirical strategy. The estimation results are reported in section 3.4. Section 3.5 provides an analysis of the bias behavior and its effect on the slope coefficient estimates. We give some concluding remarks and suggestions for further research in Section 3.6.

3.2 Empirical Tests of Uncovered Interest Rate Parity (UIP)

The uncovered interest rate parity (UIP) is an equilibrium condition stating that the expected return on a domestic asset denominated in domestic currency should equal the expected return on a foreign asset denominated in foreign currency, if they only differ with respect to the currency of denomination.[1] If i_t is the interest rate on the domestic asset between time t and $t+1$, i_t^* is the interest rate on the foreign asset between time t and $t+1$, s_t

[1] In this sense, the domestic and the foreign assets should be either risk-free or equally risky. Therefore, the only relevant source of risk is the nominal exchange rate.

is the spot exchange rate (the price of foreign currency in units of domestic currency) and $E_t(s_{t+1})$ is the expectation (at time t) for the future value of the spot exchange rate at time $t+1$, then the uncovered interest rate parity condition can be expressed by the following equation:

$$E_t\Big(\frac{s_{t+1}}{s_t}\Big)(1+i_t^*) = (1+i_t) \qquad (3.1)$$

Since the market expectation for the future value of the spot exchange rate $E_t(s_{t+1})$ is not directly observable, empirical studies usually replace it with the future realization of the exchange rate at time $t+1$.[2] In this way, they jointly test the UIP with the rational expectations hypothesis, which states that future realizations equal current rational expectations plus a white noise error-term. For reasonably small interest rates and changes in the exchange rate, condition (3.1) can be approximated as:

$$\Delta s_{t+1} + i_t^* \approx i_t \qquad (3.2)$$

Where Δs_{t+1} is the realized change of the exchange rate (in percentage terms) between time t and $t+1$.

3.2.1 Conventional Empirical Tests of UIP

Usually, empirical tests of UIP are not based on equation (3.2), but on a modified version of it. In fact, since it is widely believed that interest rates follow unit-root, or near-unit-root processes, the foreign interest rate is subtracted from both sides of equation (3.2) in order to get (approximately) stationary processes. Therefore, empirical tests are normally conducted by regressing the change in the exchange rate on the interest rate differential according to the following specification:

$$\Delta s_{t+1} = \alpha + \beta_1(i_t - i_t^*) + \epsilon_{1,t+1} \qquad (3.3)$$

[2] The exception are studies that use survey data on exchange rate expectations, see Frankel and Froot (1987), Chinn and Frankel (2002) or Chinn (2006), for example.

where α is a constant, possibly time-invariant risk premium, and $\epsilon_{1,t+1} \sim iid(0, \sigma^2_{\epsilon_1})$. If UIP holds, then the slope coefficient β_1 should not differ significantly from 1. Therefore, conventional empirical tests of UIP are tests of the null hypothesis that $\beta_1 = 1$.[3]

3.2.2 Evidence from Conventional Tests

Most empirical studies that employ regressions like (3.3) in order to test for UIP find that the interest rate differential fails to explain subsequent changes in the nominal exchange rate. Moreover, not only is the slope coefficient significantly different from 1, but it is usually significantly negative. In fact, negative values for the slope coefficient estimates are documented in numerous surveys of the empirical literature on UIP (see for example Froot and Thaler, 1990; MacDonald and Taylor, 1992; Engel, 1996). Values as low as −3 are found to be quite common when UIP is tested using the conventional regressions. Finally, recent evidence using short term data (between 3 months and 12 months horizons) is not more favorable to the UIP either (Chinn and Meredith, 2005; Chinn, 2006).

Before formally introducing the dataset and presenting the formal empirical specification in section 3.3, here we provide some illustrative evidence about the slope coefficient values obtained in these conventional or difference regressions. As an example, we consider difference regressions of 9 industrial countries against the US (dollar).[4]

Results from these difference regressions (conventional tests) are reported in tables 3.1 and 3.2. In line with most of these empirical studies, the slope coefficient estimates are negative in all, but one case (Italy). In order to further investigate the UIP proposition, we test the null hypothesis that the slope coefficient equals unity. The results, reported in the lower panel of this table, indicate that the null hypothesis can be rejected in 6 out of the 9 cases

[3]In addition, some authors also check whether there is a foreign exchange premium by testing the null hypothesis $\alpha = 0$.

[4]The choice of US as an anchor country in this example is made for illustrative purposes only. Most conclusions hold when any of the other 9 countries in our sample is used as an anchor, see section 3.4.

(countries) at the 5 percent significance level. At the 10 percent significance level, the null hypothesis is rejected in all but one case (Italy).

Furthermore, the panel data estimations of the difference regression specification produce similar results as the time series estimations. Table 3.2 reports negative slope coefficient estimates for the random-effects and the fixed-effects estimations. Moreover, the fixed-effects estimate is significantly negative even at 1 percent significance level. Finally, the lower panel of Table 3.2 shows that the null hypothesis that the slope coefficient equals unity can be rejected at any conventional significance level (e.g. 1 percent significance). In conclusion, there is substantial evidence that the slope coefficient(s) estimated with the conventional regression specification significantly differ(s) from unity and is generally negative.

Table 3.1: Conventional Regressions

	BEL	CAN	FRA	GER	ITA	JAP	NL	SWI	UK
$(i - i^*)$	-0.163	-1.177	-0.028	-0.776	0.442	-2.759	-1.765	-1.159	-1.389
	(-0.27)	(-2.13)**	(-0.05)	(-1.14)	(0.90)	(-3.30)***	(-2.60)***	(-1.80)*	(-2.03)**
Constant	0	0.001	0.001	-0.002	0.001	-0.009	-0.002	-0.005	0.004
	(-0.03)	(1.49)	(0.37)	(-1.07)	(0.25)	(-3.01)***	(-1.35)	(-1.93)*	(1.74)*
Obs	333	374	374	374	333	332	374	374	374
R^2	0	0.01	0	0	0	0.03	0.02	0.01	0.01
Test: $\beta = 1$									
F-value	F(1,331)=3.69	F(1,372)=15.57	F(1,372)=3.41	F(1,372)=6.81	F(1,331)=1.28	F(1,330)=20.26	F(1,372)=16.58	F(1,372)=11.22	F(1,372)=12.16
p-value	0.0558	0.0001	0.0658	0.0094	0.2587	0.000	0.0001	0.0009	0.0005

Note: The table presents results from time-series estimations of the difference (conventional) regression equation for testing the uncovered interest rate parity (UIP) condition given by: $\Delta s_{t+1} = \alpha + \beta(i_t - i_t^*) + \epsilon_{t+1}$. The nominal exchange rates s_t are defined as number of (domestic) currency units per US dollar.i_t^* refers to the one-month Eurocurrency deposit interest rate for the US dollar, while i_t refers to the corresponding rate for each partner country. t-statistics, calculated using Newey-West consistent standard errors, are reported in parentheses in the row below the test statistics. Significance at the 10%, 5%, and 1% significance level is denoted by *, **, and ***, respectively. The lower panel in the table contains results from F-tests based on the null hypothesis that the slope coefficient is unity ($\beta = 1$). The first row contains F-test statistics based on 330-372 degrees of freedom, while the second row reports the corresponding p-values.

Table 3.2: Conventional (Difference) Tests: Panel Estimation Results

	RE	FE
interest differential	-0.283	-0.732
	(-1.67)*	(3.50)***
Constant	0	0
	(-0.28)	(-0.36)
Observations	3242	3242
Number of code	9	9
R-squared		0
Test: $\beta = 1$		
F-value	F(1,3240)=57.13	F(1,3232)=68.64
p-value	0.0000	0.0000

Note: The table presents results from panel data estimations of the difference (conventional) regression equation for testing the uncovered interest rate parity (UIP) condition given by: $\Delta s_{t+1} = \alpha + \beta(i_t - i_t^*) + \epsilon_{t+1}$. The nominal exchange rates s_t are defined as number of (domestic) currency units per US dollar.i_t^* refers to the one-month Eurocurrency deposit interest rate for the US dollar, while i_t refers to the corresponding rate for each partner country. We report random-effects (RE) and fixed-effects (FE) estimation results, respectively. t-statistics, calculated using Newey-West consistent standard errors, are reported in parentheses in the row below the test statistics. Significance at the 10%, 5%, and 1% significance level is denoted by *, **, and ***, respectively. The lower panel in the table contains results from F-tests based on the null hypothesis that the slope coefficient is unity ($\beta = 1$).The first row contains F-test statistics based on 3232-3240 degrees of freedom, while the second row reports the corresponding p-values.

3.2.3 Comparison: Results from Level and Difference Regressions

An alternative UIP test would be to regress the domestic currency holding period return of a deposit denominated in foreign currency (which equals the foreign interest rate plus the change in the nominal exchange rate) on the holding period return of a domestic deposit (domestic interest rate).[5] Here we present results from such "level regressions" for each of the other 9 countries against the USA and contrast them with the estimation results from the difference (conventional) regressions reported before. For a similar, level specification used in testing for UIP in the inter-war period (gold exchange standard) see Bordo and MacDonald (2003), for example.

The level results are presented in tables 3.3 and 3.4. Table 3.3 contains results from separate time-series estimations per country relative to the US (dollar). Contrary to the results from the difference regressions, the estimates for the slope coefficient reported in Table 3.3 are generally positive (with the exception of the estimates for Japan and Switzerland) and reasonably close to the theoretically expected value of unity. Note that point estimates are found

[5] This regression specification implies that the variables are either stationary or cointegrated. Section 3.3.4 presents stationarity tests.

to be both above and below unity. In order to further investigate this issue, we test the null hypothesis that the slope coefficient equals unity. The test results suggest that the null hypothesis can be rejected only for one country (Japan) at the 10 percent significance level.

Table 3.4 contains results from the panel data estimations of the level regression specification. Both random-effects and fixed-effects coefficient estimates are below, but very close to the theoretically expected value of unity. In fact, the lower panel of this table contains results from tests of the null hypothesis that the true slope coefficient equals unity. The F-values of these tests are low (and correspondingly, the p-values are high), suggesting that the null hypothesis cannot be rejected at any reasonable significance level (e.g. at 10 percent significance). Finally, the estimated values of the slope coefficient estimates deserve special attention: they summarize the "average" relationship between holding period returns in US dollars and holding period returns in all other 9 currencies. Overall, the results from these level regressions provide suggestive evidence for a significantly positive slope coefficient in a region around unity.

Table 3.3: Level Regressions

	BEL	CAN	FRA	GER	ITA	JAP	NL	SWI	UK
interest rate (i)	1.306	0.871	1.158	1.545	1.078	-0.293	0.389	-0.179	0.608
	(2.67)***	(3.57)***	(2.98)***	(1.93)*	(3.39)***	(-0.41)	(0.55)	(-0.21)	(1.23)
Constant	-0.002	0	-0.002	-0.002	-0.002	0.005	0.003	0.004	0.002
	(-0.49)	(-0.28)	(-0.54)	(-0.53)	(-0.55)	(1.81)*	(0.8)	(1.41)	(0.45)
Observations	333	374	374	374	333	332	374	374	374
R-squared	0.02	0.03	0.02	0.01	0.03	0	0	0	0
Test: $\beta = 1$									
F-value	F(1,331)=0.39	F(1,372)=0.28	F(1,372)=0.17	F(1,372)=0.46	F(1,331)=0.06	F(1,330)=3.24	F(1,372)=0.75	F(1,372)=1.88	F(1,372)=0.63
p-value	0.533	0.597	0.684	0.497	0.807	0.073	0.387	0.171	0.437

Note: The table presents results from time-series estimations of the level (unconventional) regression equation for testing the uncovered interest rate parity (UIP) condition given by: $\Delta s_{t+1} + i_t^* = \alpha + \beta i_t + \epsilon_{t+1}$. The nominal exchange rates s_t are defined as number of (domestic) currency units per US dollar.i_t^* refers to the one-month Eurocurrency deposit interest rate for the US dollar, while i_t refers to the corresponding rate for each partner country. t-statistics, calculated using Newey-West consistent standard errors, are reported in parentheses in the row below the test statistics. Significance at the 10%, 5%, and 1% significance level is denoted by *, **, and ***, respectively. The lower panel in the table contains results from F-tests based on the null hypothesis that the slope coefficient is unity ($\beta = 1$). The first row contains F-test statistics based on 330-372 degrees of freedom, while the second row reports the corresponding p-values.

Table 3.4: Level Regressions: Panel Estimation Results

	RE	FE
interest rate (i)	0.81	0.88
	(5.68)***	(5.38)***
Constant	0.001	0.001
	(-1.05)	(-0.59)
Observations	3242	3242
Number of code	9	9
R-squared		0.01
Test: $\beta = 1$		
F-value	F(1,3240)=1.78	F(1,3232)=0.54
p-value	0.1823	0.4642

Note: The table presents results from panel data estimations of the level (unconventional) regression equation for testing the uncovered interest rate parity (UIP) condition given by: $\Delta s_{t+1} + i_t^* = \alpha + \beta i_t + \epsilon_{t+1}$. The nominal exchange rates s_t are defined as number of (domestic) currency units per US dollar.i_t^* refers to the one-month Eurocurrency deposit interest rate for the US dollar, while i_t refers to the corresponding rate for each partner country. Random-effects (RE) and fixed-effects (FE) estimation results are reported. t-statistics, calculated using Newey-West consistent standard errors, are reported in parentheses in the row below the test statistics. Significance at the 10%, 5%, and 1% significance level is denoted by *, **, and ***, respectively. The lower panel in the table contains results from F-tests based on the null hypothesis that the slope coefficient is unity ($\beta = 1$).The first row contains F-test statistics based on 2900-3240 degrees of freedom, while the second row reports the corresponding p-values.

3.2.4 Biasedness of the Slope Coefficient Estimator When UIP Does Not Hold

The results from the difference and level regressions suggested that the empirical evidence about the validity of UIP crucially depends on the exact regression specification used in the estimations. In this section, we explore an alternative reason why slope coefficient estimates from conventional (difference) empirical tests of UIP might be biased. We base our argument on the possibility (as suggested by the level regressions) that the "true" slope coefficient might very close, but different from the theoretically-expected value of unity.

Equation (3.3), used in conventional empirical tests of UIP, is derived after subtraction of i_t^* from both sides of equation (3.2), and after a specific type of parametrization. If the true data-generating process is not exactly described by equation (3.2) (with a slope coefficient of one), then the expected value of the estimator $\hat{\beta}_1$ from equation (3.3) might not correspond to the "true" slope coefficient.[6]

In fact, let us assume that the true data-generating process can be approx-

[6]This demonstration closely follows the approach taken in Kool and Thornton (2004) and Thornton (2006) for tests of the expectations hypothesis.

imated by the following equation:[7]

$$\Delta s_{t+1} + i_t^* = \beta i_t + \epsilon_{t+1} \tag{3.4}$$

where the slope coefficient β does not necessarily equal 1. There are several reasons why the slope coefficient from the "true" data-generating process, i.e. the "true β" from equation (3.4), might differ from unity. First, this condition should hold with unity only in expectation, and not necessarily in realization. Since market expectations of future changes in the nominal exchange rate are not directly observable ex-ante (at the time of their formation), they are approximated by the ex-post realized changes in the nominal exchange rate. Therefore, the "true β" in the process described by equation (3.4) will generally differ from unity even in the absence of any additional imperfections (in general, any deviations from the underlying assumptions of the UIP model).[8] Second, even disregarding the above argument based on the empirical procedure and the difficulty in measuring the "true" market expectations ex-ante, the slope coefficient might (slightly) differ from the theoretically expected value of unity because of several market distortions or imperfections. For example, there might be transaction costs, risk premium considerations (including the peso-problem), or differences in market participants' expectations as they might be based on (slightly) different information sets.[9]

Then, subtracting i_t^* from both sides of equation (3.4) and slightly rearranging leads to:

$$\Delta s_{t+1} = \beta(i_t - i_t^*) + (\beta - 1)i_t^* + \omega_{t+1} \tag{3.5}$$

[7]This equation corresponds to the equalization of the holding period returns on domestic and foreign assets. For alternative tests of UIP using this regression specification see Lothian and Wu (2005). Note that this cannot literally be the true specification. Reversing the domestic and foreign country in equation (3.4) immediately makes clear that both equations cannot simultaneously hold unless "true" $\beta = 1$.

[8]This is true to the extent that the ex-post realizations do not perfectly correspond with the ex-ante market expectations at every point in time.

[9]The list of reasons can be extended with the inclusion of different manifestations of irrational behavior.

The above equation only exactly corresponds to equation (3.3) that is usually estimated in empirical tests of UIP when $\beta = 1$. If not, the expected value of the least-squares estimator for the slope coefficient from equation (3.3) will generally differ from the "true" slope coefficient:

$$E\hat{\beta} = \beta + (\beta - 1)E\frac{\Sigma(\bar{i}_t - \bar{i}_t^*)\bar{i}_t^*}{\Sigma(\bar{i}_t - \bar{i}_t^*)^2} - E\frac{\Sigma(\bar{i}_t - \bar{i}_t^*)\bar{\omega}_t}{\Sigma(\bar{i}_t - \bar{i}_t^*)^2} \tag{3.6}$$

where the "bar" denotes variables adjusted for the mean. Or, expressed in terms of variances and covariances:

$$\begin{aligned} E\hat{\beta} &= \beta + (\beta - 1)\left[\frac{Cov(i - i^*, i^*)}{Var(i - i^*)}\right] + \left[\frac{Cov(i - i^*, \omega)}{Var(i - i^*)}\right] = \\ &= \beta + (\beta - 1)\left[\frac{Cov(i, i^*) - Var(i^*)}{Var(i) - 2Cov(i, i^*) + Var(i^*)}\right] + \\ &+ \left[\frac{Cov(i, \omega) - Cov(i^*, \omega)}{Var(i) - 2Cov(i, i^*) + Var(i^*)}\right] \end{aligned} \tag{3.7}$$

The probability limit of this expression is given by:

$$P\lim_{N\to\infty}\hat{\beta} = \beta + (\beta - 1)\left[\frac{Cov(i, i^*) - Var(i^*)}{Var(i) - 2Cov(i, i^*) + Var(i^*)}\right] \tag{3.8}$$

Therefore, as long as $\beta \neq 1$, the second term in equation (3.8) will differ from zero, and $\hat{\beta}$ from conventional tests will be a biased estimate of the true slope coefficient. The relationship between the interest rate distribution(s) and the total bias (given by the bracketed expression in equation (3.8)) can be better seen by a slight rearrangement of equation (3.8):

$$P\lim_{N\to\infty}\hat{\beta} = \beta + (\beta - 1)\left[\frac{\rho\delta^{1/2} - \delta}{1 - 2\rho\delta^{1/2} + \delta}\right] \tag{3.9}$$

where ρ is the coefficient of correlation between domestic and foreign interest rates, and δ is the ratio of their variances defined as $\delta = \frac{Var(i^*)}{Var(i)}$. If the foreign (anchor country) interest rate is more volatile than the domestic interest rate, i.e., if $\delta > 1$, the expression for total bias, given between square brackets, will be strictly negative.[10] In this case, if $\beta < 1$, then relatively high foreign interest rate volatility will translate into an upward-biased estimate $\hat{\beta}$ of the slope coefficient. For $\delta < 1$, the bias term can be both positive and negative.

For an analysis of the relations between ρ, δ, and bias we start with equation (3.9), which implies that the (average) slope coefficient estimates should be given by:[11]

$$\hat{\beta} = \beta + (\beta - 1)Bias \tag{3.10}$$

If we assume that the "true" slope coefficient β is in the range (0.5-1.5),[12] then equation (3.10) suggests that the total bias expression should be either higher than 1 or lower than -3 in order to explain the negative slope coefficient estimates found in the estimations.[13] In order to understand whether the bias can theoretically explain the values for the slope coefficient found in the estimations, we graphically analyze the behavior of the total bias expression with respect to the main parameters (ρ and δ).[14]

Figure 3.1 describes the behavior of the total bias expression for all admissible combinations of the parameters ρ and δ.[15] It contains a three-dimensional plot viewed from two different angles and presents values for the total bias expression as a complex function of ρ and δ. Moreover, the plot shows the precise location of low and high value areas for the bias expression within the

[10] Kool and Thornton (2004) focus on the case where $\delta > 1$ because short-term interest rates are typically more volatile than long-term rates. The bias is strictly negative in this case.

[11] Equation (3.9) will not hold if there are additional distortions: time-varying risk-premia, additional sources of (estimation) bias, misspecifications, etc.

[12] This range of values corresponds with the slope coefficient estimates from the level regressions reported in Tables 3.3 and 3.4.

[13] In fact, the condition that the total bias expression should be either higher than 1 or lower than -3 is necessary to explain slope coefficient estimates below zero. In order to explain the average slope coefficient estimate found in the estimations (-2), the total bias should either get much higher values than 1 or much lower than -3.

[14] Here we provide a graphical analysis of bias behavior for all possible values of ρ ($-1 \leq \rho \leq 1$). Since bias behavior changes abruptly in the vicinity of $\rho = 1$, compared to $\rho = 0.9$ for example, we present the analysis for $\rho \leq 0.9$ in the appendix.

[15] The values for δ in the figures are restricted only for illustrative purposes. Similar figures for alternative value intervals for δ ($\delta < 100$) are available upon request.

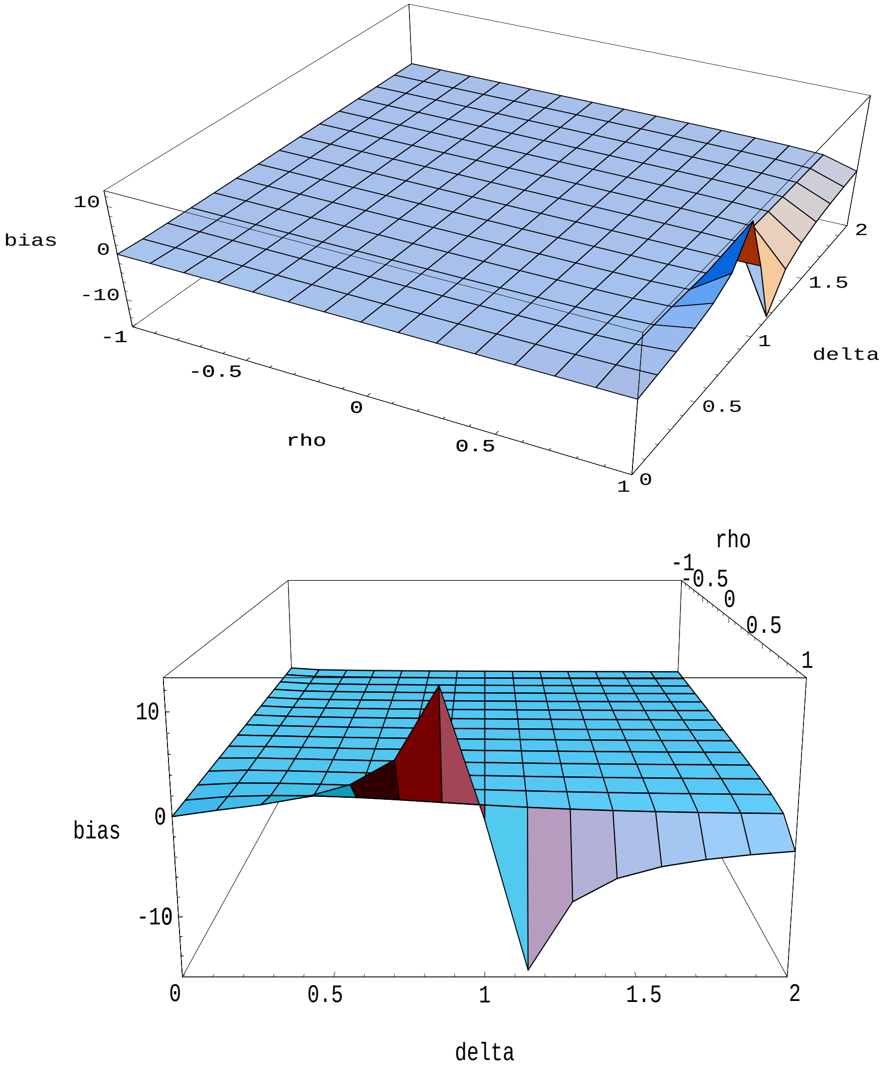


Figure 3.1: Total Bias Values for Alternative Combinations of δ and ρ (3D Projections for $\rho \leq 1$)

Note: The three-dimensional plots present values for the total bias expression for alternative combinations of δ (relative interest rate volatility) and ρ (coefficient of correlation between the corresponding interest rate series). The relative volatility indicator (delta) can take any value in the interval $(0, 2)$, while the coefficient of correlation can take any value in the interval $(-1, 1)$. For convenience and clarity in presentation, we restrict the values for the relative interest rate volatility ($0 \leq \delta \leq 2$). In extremis, the total bias expression is not defined for $\rho = 1$ and $\delta = 1$.

ρ-δ plane: combinations of very high ρ with δ close but lower than one ($\delta < 1$) produce extremely high values, while combinations of very high ρ with δ close but higher than one ($\delta > 1$) produce extremely low values. In fact, this finding suggests that the total bias expression gets large values (in absolute terms) for very restrictive combinations of ρ and δ. Only when ρ is at least about 0.95, while δ is close to unity (for example $0.5 < \delta < 1.5$), the total bias gets large enough values that are needed to explain the (negative) values for the slope coefficient estimates according to equation (3.10).[16]

3.3 Data Description and Empirical Strategy

3.3.1 Empirical Strategy

The empirical analysis proceeds as follows. First, we briefly describe the dataset used in this study (section 3.3.2, and then present summary statistics (section 3.3.3) and stationarity test for the series employed in the empirical analysis (section 3.3.4). In section 3.4 we estimate several regression specifications. First, we include an omitted variable into the conventional specification according to equation (3.5) and compute implied values for the "true" slope coefficient (section 3.4.1). Then, we investigate the relevance of the biasedness hypothesis elaborated in section 3.2.4 for the results from conventional empirical tests of UIP. For this purpose, we compare results from conventional tests obtained from specifications that use alternative anchor countries.

3.3.2 Dataset

The dataset consists of monthly observations for the following 10 industrial countries: Belgium, Canada, France, Germany, Italy, Japan, Netherlands, Switzerland, UK and USA over the period January 1975-December 2004.[17] The monthly rates of change for the nominal exchange rates and the interest rate series for one-month Eurocurrency deposits are retrieved from Datastream.[18]

[16] In fact, its maximal value is above 10, while its minimum gets close to -15.

[17] The interest rate time-series only start in 1978 for Belgium, Italy and Japan.

[18] We use interest rates on one-month Eurocurrency deposits in order to focus on currency risk only and refrain from any additional types of risk.

3.3.3 Summary Statistics

Summary statistics for all data-series employed in the empirical analysis are given in Table 3.5. All figures are annualized and expressed in percentage terms. The first two columns report the means and standard deviations for the interest rate series, while the last two columns report the corresponding statistics for the nominal exchange rate changes.

There are four interesting observations in this table. First, the interest rate series display large differences across countries (currencies) in terms of average values as well as average volatility over the entire period.[19] For example, the average interest rate for Italy (highest in the sample) is almost three times larger than the average interest rate for Switzerland (lowest in the sample). Moreover, similar differences emerge in terms of average volatility - the values for the annualized standard deviation range from slightly higher than 0.7 percent for Germany and Switzerland to almost 1.9 percent for Italy. Second, higher interest rate currencies on average depreciate against the US, while lower interest rate currencies appreciate. This provides casual evidence in support of UIP. Third, in contrast to the evidence for interest rates, the nominal exchange rate changes display similar volatility across different countries.[20] Fourth, the nominal exchange rate changes show much higher average volatility, but also much lower average values than the interest rate series.[21] In sum, the observations in Table 3.5 suggest that most differences across countries stem from the behavior of the interest rate series and not from the exchange rate changes.

[19]Throughout the empirical analysis, the terms domestic (foreign) country and domestic (foreign) currency are used interchangeably.

[20]The last column in Table 3.5 shows that the annualized standard deviation for the nominal exchange rate changes is typically about 10 percent. The only exception is the Canadian dollar as its variability vis-*à*-vis the US dollar is about half the typical figure for the other countries.

[21]In fact, the average standard deviation for the nominal exchange rate changes (10.56 percent) is about ten times larger than the corresponding figure for the interest rate series (1.12 percent), while the average change in the nominal exchange rate (0.29 percent) is much lower than the average interest rate for the whole sample (6.72 percent).

Table 3.5: Summary Statistics: Interest Rates and Exchange Rate Changes

	interest rate		exchange rate	
	mean	st.dev	mean	st.dev
BEL	6.83	1.31	-0.25	11.25
CAN	7.71	1.14	0.54	5.29
FRA	8.22	1.45	0.64	10.82
GER	5.22	0.72	-1.28	11.08
ITA	10.48	0.88	2.89	10.77
NL	3.54	0.94	-2.99	11.92
JAP	5.72	0.82	-1.02	11.01
SWI	3.51	0.74	-2.11	12.29
UK	9.18	1.12	0.96	10.57
USA	6.78	1.06		
mean	6.72	1.12	-0.29	10.56

Note: The table contains annualized summary statistics (means and standard deviations) for interest rates and exchange rate changes for 10 industrialized countries over the time period January 1975-December 2004. Due to data limitations, the summary statistics for the interest rate series for Belgium, Italy, and Japan are calculated over the period July 1978-December 2004. All data-series are retrieved from Datastream. The statistics for the interest rate series are calculated using monthly observations on the one-month Eurocurrency deposit rate for the corresponding country. The statistics for the exchange rate are calculated using monthly observations (end-of-month observations) on the nominal exchange rate against the US dollar (units of currency per US dollar) for the corresponding currency/country. All figures presented in the table are annualized and expressed in percentage terms (rounded to two decimal places).

3.3.4 Stationarity Tests

Section 3.2.3 demonstrated that the evidence from the level regressions presented in tables 3.3 and 3.4 gives some support for the validity of the UIP condition. However, a standard objection against such specification concerns the (assumed) nonstationarity of the (domestic) interest rate, which is used as an independent variable in these specifications. In order to capture the severity of the nonstationarity problem, we conduct unit-root tests for all series used in sections 3.2.2 and 3.2.3 (both level regressions and difference regressions).

We conduct two types of stationarity tests over the period January 1975-December 2004: time series unit-root tests per individual country and panel unit-root tests including observations on all countries in the dataset. For the individual country results, we employ the Augmented Dickey-Fuller (ADF) tests with drift parameter.[22] The panel unit-root tests are based upon the approach proposed by Maddala and Wu (1999). For each time-series test, we determine the optimal lag-length according to the Akaike Information Criterion (AIC).[23] Table 3.6 contains the results from stationarity (unit-root) tests.

The results in the table give strong evidence against the null hypothesis of nonstationarity (presence of unit-root). Allowing for a drift parameter in the testing specification, the nonstationarity null hypothesis can be rejected in all cases in the first three panels (Panels A, B, and C) even at the 1 percent significance level. In line with previous studies, these results imply that both variables used in the conventional tests of UIP as well as the dependent variable used in level regressions do not contain unit-root(s).

The last panel (Panel D) presents results for the domestic interest rate, which is used as explanatory (independent) variable in the level regressions. Nonstationarity can be rejected in 8 out of 9 time-series at least at the 5 percent significance level. Though the evidence against nonstationarity in Panel D is somewhat weaker compared to the first three panels, it still suggests that most (domestic) interest rate series are (close to) stationary.[24] The results from the Maddala-Wu panel unit-root tests, reported in the last column of the table, strengthen the evidence against nonstationarity: the null hypothesis of panel unit-root can be rejected at the 0.1 percent significance level for the first three series, and at the 5 percent for the last series. In sum, we conclude that nonstationarity is not a (very) serious issue for these variables, and therefore, most estimation results in Tables 3.1 and 3.4 can be considered valid.

[22]We also tested for (non)stationarity using the unit-root testing procedure employed by Kwiatkowski, Phillips, Schmidt, and Shin (1992). Generally, the results from these KPSS tests are even more in favor of stationarity. Results from these tests are available upon request.

[23]The use of alternative information criteria (for example (SBIC) or (HQIC)) in the selection of the optimal lag-length does not lead to significantly different results and conclusions. The panel-unit root tests (Maddala-Wu test statistics) are based on the optimal lag-orders of the corresponding series, also determined according to AIC. We only report results for the Maddala-Wu panel unit-root tests with drift. The complete results from alternative specifications (without trend, drift, excluding certain countries/series, including alternative lag-orders) are available in the tables in the appendix (section A.1).

[24]The evidence of nonstationarity in the interest rate series is limited to certain countries: Japan at the optimal lag-length and Italy at most higher lag-lengths (Table A.4).

Table 3.6: Stationarity Tests

Country	BEL	CAN	FRA	GER	ITA	JAP	NL	SWI	UK	ALL (Panel)
Panel A: Change in the Nominal Exchange Rate										
Test statistic	-10.03***	-19.69***	-18.96***	-18.99***	-9.83***	-19.04***	-10.17***	-18.76***	-17.99***	90.00****
Lag-order (AIC)	2	0	0	0	2	0	2	0	0	
Panel B: Interest Rate Differential										
Test statistic	-3.39***	-3.91***	-4.63***	-2.35***	-3.51***	-3.62***	-3.71***	-2.72***	-4.32***	66.21****
Lag-order (AIC)	3	2	2	3	4	2	3	1	1	
Panel C: Domestic Currency Holding Period Returns of US Dollar Deposits										
Test statistic	-9.45***	-18.38***	-9.35***	-9.73***	-9.29***	-18.68***	-9.65***	-10.09***	-12.38***	90.00****
Lag-order (AIC)	2	0	2	2	2	0	2	2	1	
Panel D: Interest Rates (Returns) in Domestic Currency										
Test statistic	-1.87**	-1.89**	-2.90***	-1.82**	-2.12**	-1.21	-2.68***	-2.22**	-2.13**	31.91**
Lag-order (AIC)	3	3	2	4	2	2	4	4	4	

Note: The table reports Augmented Dickey-Fuller (ADF) and and Maddala-Wu (MW) test statistics based on the null hypothesis that the series follows a unit-root process. Due to limitations in the interest rate data, the tests for Belgium, Italy, and Japan are conducted over the period July 1978-December 2004. For all other countries, the tests are conducted over the entire period January 1975-December 2004. Panels A and B contain results for series used in conventional (difference) tests of UIP reported in tables 3.1 and 3.2. Panel A refers to nominal exchange rate changes relative to the US dollar, while Panel B refers to the interest rate differential. Panels C and D contain results for series used in level tests of UIP reported in tables 3.3 and 3.4. Panel C refers to domestic currency holding period return on US dollar denominated deposits (dependent variable in level regressions), while Panel D refers to the (domestic country) interest rate (independent variable in the level regressions). The optimal lag-order (lag-length) is selected according to the Akaike Information Criterion (AIC). The critical ADF values (allowing for a drift parameter) with 316-332 degrees of freedom are given as follows: -1.284 (10% significance level), -1.649 (5%), -2.337 (1%). The last column reports panel unit-root test statistics (with drift) calculated according to Maddala and Wu (1999). The Maddala and Wu (1999) panel unit-root test is a one sided test that follows χ^2-distribution with $2n$ degrees of freedom (where n is the number of separate time-series). The (upper) critical values for the χ^2-distribution with 18 degrees of freedom (tests for the entire dataset) are given as follows: 25.989 (10% significance level), 28.869 (5%), 34.805 (1%), 42.312 (0.1%). All test statistics are rounded to two decimal places. Rejection of the null hypothesis at 10%, 5% and 1% significance level is indicated by *, **, and ***, respectively. For the Maddala-Wu panel unit-root tests, the null hypothesis of nonstationarity (presence of panel unit-root) is rejected in each case at least at the 10% significance level. Moreover, significance at 10%, 5%, 1%, and 0.1% is indicated by *, **, ***, and ****, respectively.

3.4 Estimation Results

3.4.1 Omitted Variable Bias

If the "true β" from the true data-generating process is close to, but different from 1, then equation (3.5) suggests that the anchor country interest rate i^* will act as an omitted variable in the conventional UIP specification. Hence, we can directly test for the importance of this omitted variable bias by including i^* as an additional independent variable in the regression equation. Therefore, we estimate the following specification:

$$\Delta s_{t+1} = \alpha_0 + \beta_{(i-i^*)}(i_t - i_t^*) + \beta_{(i^*)} i_t^* + \omega_{t+1} \tag{3.11}$$

If this omitted variable is important, then its slope coefficient $\beta_{(i^*)}$ should significantly differ from zero. We treat each of the 10 countries as anchors, and run regressions against the remaining 9 partner (domestic) countries. The results from the corresponding panel estimations are reported in Table 3.7 and suggest that the anchor country interest rate i^* is an important variable missing in the conventional tests of UIP. In fact, its effect is significant for 7 out of 10 countries at the 5 percent significance level, and for 8 out of 10 countries at the 10 percent significance level.

Although the omitted variable is significant in most cases, its inclusion does not seem to provide additional support for UIP. In fact, the (unrestricted) coefficients in front of the interest rate differential either stay negative or turn only marginally positive. Only for Germany (random-effects estimation), it is significantly positive. We conclude that the omission of the anchor country interest rate might lead to misspecification of the regression equation used in conventional UIP, but it does not explain the UIP puzzle.

Table 3.7: UIP Specification Including Omitted Variable i^* (Panel Estimations)

Anchor Country	BEL	CAN	FRA	GER	ITA	JAP	NL	SWI	UK	USA
					Random-Effects Panel Estimations					
i-i*	0.126	-0.295	0.210	0.219	-0.246	-0.380	-0.085	-0.033	-0.472	-0.088
	(0.93)	(-1.51)	(1.48)	(1.73)*	(-1.33)	(-1.76)*	(-0.68)	(-0.24)	(-2.89)***	(-0.50)
i*	-0.252	0.300	-0.401	0.079	-0.501	1.301	0.657	0.920	0.026	0.786
	(-2.1)**	(1.74)*	(-3.35)***	(0.43)	(-3.67)***	(5.77)***	(4.03)***	(4.47)***	(0.17)	4.35
Constant	0.001	-0.003	0.002	0.000	0.001	-0.001	-0.002	-0.001	-0.002	-0.005
	(1.60)	(-2.24)**	(2.63)***	(0.28)	(1.32)	(-1.10)	(-2.71)***	(-1.22)	(-1.73)*	(-3.97)***
					Fixed-Effects Panel Estimations					
i-i*	-0.078	-0.952	-0.170	0.018	-0.637	-0.769	-0.469	-0.257	-1.053	-0.474
	(-0.46)	(-3.69)***	(-0.86)	(0.11)	(-2.73)***	(-2.94)***	(-3.00)***	(-1.59)	(-5.24)***	(-2.15)***
i*	-0.328	0.146	-0.600	0.113	-0.734	1.287	0.670	0.916	-0.131	0.683
	(-2.60)***	(0.83)	(-4.30)***	(0.61)	(-4.57)***	(5.70)***	(4.11)***	(4.45)***	(-0.82)	(3.71)***
Constant	0.002	-0.002	0.003	0.000	0.002	-0.000	-0.002	-0.001	-0.002	-0.004
	(2.05)**	(-1.92)*	(3.41)***	(0.42)	(1.93)*	(-0.04)	(-2.37)**	(-0.44)	(-1.82)*	(-3.45)***
Observations	2996	3242	2968	3242	2719	2988	3242	3242	3242	3242
Number of groups										
R-squared	0.003	0.006	0.009	0	0.009	0.014	0.008	0.007	0.009	0.008

Note: The table presents results from panel data estimations of the difference (conventional) regression equation for testing the uncovered interest rate parity (UIP) condition, which includes the (anchor country) interest rate i_t^* as an omitted variable. The empirical specification estimated is given as follows: $\Delta s_{t+1} = \alpha + \beta_{(i-i^*)}(i_t - i_t^*) + \beta_{(i^*)} i_t^* + \omega_{t+1}$. The nominal exchange rates s_t are defined as number of (domestic) currency units per unit of anchor country currency. i_t^* refers to the one-month Eurocurrency deposit interest rate in the anchor country currency, while i_t refers to the corresponding rate for each partner (domestic) currency/country. The upper panel reports estimation results from the random-effects procedure, while the lower panel reports estimation results following the fixed-effects procedure. t-statistics, calculated using Newey-West consistent standard errors, are reported in parentheses in the row below the test statistics. Significance at the 10%, 5%, and 1% significance level is denoted by *, **, and ***, respectively.

Table 3.8: Implied Values for True β

anchor country	BEL	CAN	FRA	GER	ITA	JAP	NL	SWI	UK	USA
Implied Values for True β										
implied β (RE)	0.436	0.502	0.404	0.649	0.126	0.960	0.786	0.943	0.277	0.848
implied β (FE)	0.296	0.097	0.115	0.565	-0.185	0.758	0.601	0.829	-0.091	0.604

Note: The table presents implied values for the "true" slope coefficient β calculated using the panel data estimation results from Table 3.7 according to equation (3.12): $\beta = \frac{\hat{\beta}_{(i-i^*)}+\hat{\beta}_{(i^*)}+1}{2}$. In this formula $\hat{\beta}_{(i-i^*)}$ refers to the slope coefficient estimate in front of the interest rate differential term, while $\hat{\beta}_{(i^*)}$ refers to the slope coefficient estimate in front of the (anchor country) US interest rate term in equation (3.11). The first row reports values calculated using the random-effects estimates, while the second row reports values based on the fixed-effects estimates.

Restricting the Coefficients

However, from equation (3.5) we know that the coefficient in front of the interest rate difference term should equal β and the coefficient in front of the "omitted" anchor country interest rate should equal $\beta - 1$. Hence, they should sum up to $\beta + \beta - 1 = 2\beta - 1$. Using this restriction, we can calculate "implied" values for the true slope coefficient β. If $\hat{\beta}_{(i-i^*)}$ stands for the slope coefficient estimate in front of the interest rate differential and $\hat{\beta}_{(i^*)}$ for the slope coefficient estimate in front of the omitted variable i^*, then the implied value for the true slope coefficient β can be obtained as follows:

$$\beta = \frac{\hat{\beta}_{(i-i^*)} + \hat{\beta}_{(i^*)} + 1}{2} \tag{3.12}$$

The implied values for the true β are given in Table 3.8. The first row reports implied values calculated using the random-effects panel estimates, while the second row reports implied values from fixed-effects panel estimates. There are several interesting findings in this table. First, all implied values in the first row (using random-effects) are positive. Second, most of these values fall in the range $0.4 - 1.0$. The only exceptions are the estimations for Italy and UK. In sum, the evidence presented in this table suggests that by controlling for the omitted variable in conventional UIP test specifications - i^*, we can "recover" positive implied values for the true slope coefficient. Moreover, these values are typically close to the theoretically-expected value of unity when calculated using the random-effects.[25]

[25] Generally, there is a "uniform" decrease of the implied values when calculated with the fixed-effects compared to the random-effects estimates. The implied values for Canada

Instead of computing implied values from the unrestricted estimation, we may directly estimate equation (3.11) in restricted form. The results from these constrained estimations are reported in Table 3.9. All slope coefficient estimates in front of the interest rate differential turn positive, while all slope coefficient estimates in front of the omitted variable i^* turn negative.[26] Looking at the complete specification used in conventional UIP tests given by equation (3.5), the first slope coefficient estimate refers to the true β, while the second refers to $1-\beta$. The null hypothesis that the slope coefficient estimate in front of the interest rate differential term equals unity can be rejected for all anchor countries except Japan. Therefore, the results from these constrained regressions further strengthen the evidence from the "level regressions" in tables 3.3 and 3.4 that the "true" slope coefficient might be close to, but smaller than unity. In fact, the results indicate that for most (anchor) countries the "true" β is around 0.5. For Japan and US it is higher, for the UK lower. Note that in the US column the "true" β-estimate is very close to that in Table 3.4 (level regression), but very different from that reported in Table 3.2 (conventional regression).[27]

and France decrease substantially when calculated using with fixed-effects estimates. The Hausman (1979) test can be rejected at 1 percent (Canada) or at 5 percent significance (France), suggesting that only fixed-effects estimates are valid for these countries.

[26] The constant term in all specifications is very low.

[27] The US column in Table 3.9 contains results from pooled-OLS estimations where the US is used as an anchor for the other 9 countries. Therefore, we may compare these results with those for all 9 countries against the US (panel estimates) in Tables 3.2 and 3.4, respectively.

Table 3.9: Estimation Results: Constrained Linear Regression (Pooled-OLS)

Anchor Country	BEL	CAN	FRA	GER	ITA	JAP	NL	SWI	UK	USA
i-i*	0.498	0.639	0.472	0.509	0.508	0.903	0.561	0.571	0.316	0.809
	(4.66)***	(4.37)***	(4.19)***	(5.14)***	(3.70)***	(5.61)***	(5.68)***	(4.84)***	(2.44)**	(5.68)***
i*	-0.501	-0.361	-0.527	-0.491	-0.491	-0.0964	-0.438	-0.428	-0.683	-0.190
	(-4.68)***	(-2.47)**	(-4.67)***	(-4.96)***	(-3.58)***	(-0.60)	(-4.43)***	(-3.63)***	(-5.26)***	(-1.33)
Constant	0.002	0.002	0.004	0.002	0.004	-0.001	0.002	0.001	0.005	0.001
	(3.75)***	(2.44)**	(4.66)***	(3.44)***	(4.59)***	(-0.78)	(3.30)***	(1.43)	(5.87)***	(1.05)

Note: The table presents pooled-OLS estimation results from constrained linear regressions based on equation (3.11) and the following constraint: $\hat{\beta}_{(i-i^*)} - \hat{\beta}_{(i^*)} = 1$. In this constraint $\hat{\beta}_{(i-i^*)}$ refers to the slope coefficient estimate in front of the interest rate differential term, while $\hat{\beta}_{(i^*)}$ refers to the slope coefficient estimate in front of the (anchor country) interest rate term in equation (3.11). The nominal exchange rates s_t are defined as number of (domestic) currency units per unit of anchor country currency .i_t^* refers to the one-month Eurocurrency deposit interest rate in the anchor country currency, while i_t refers to the corresponding rate for each (domestic) partner currency/country. Hence, $i_t - i_t^*$ refers to the interest rate differential. t-statistics are reported in parentheses in the row below the test statistics. Significance at the 10%, 5%, and 1% significance level is denoted by *, **, and ***, respectively.

3.4.2 Results from Conventional Tests of Uncovered Interest Parity

The results from alternative UIP specifications (either level regressions or including the "omitted" variable) demonstrated that the "true" slope coefficient might be positive and even close to the theoretically expected value of unity. These findings contradict the anomalous results obtained in conventional (difference) tests of UIP. In this section, we investigate the relevance of the biasedness hypothesis - elaborated on in section 3.2.4 - for the anomalous results obtained in the conventional empirical tests. For this purpose, we estimate regressions of the type employed in conventional tests of UIP given in equation (3.3) for all possible bilateral country pairs for each of the six five-year sub-periods: 1975-1979, 1980-1984, 1985-1989, 1990-1994, 1995-1999 and 2000-2004. Hence, we treat each of the 10 countries as domestic country and run regressions against each of the remaining 9 anchor (foreign) countries. In total, we estimate 520 regressions in this way and use the estimation results from these regressions in order to test for the importance of the biasedness hypothesis in the sample at hand.[28] A caveat refers to the fact that the theoretical results obtained in section 3.2.4 hold in the limit. Empirically, we use 5 year subsamples to obtain sufficient number of independent observations/estimates.

Summary statistics for the slope coefficient estimates obtained from conventional tests of UIP are provided in table 3.10. Each column displays the mean, minimum and maximum values and the standard deviation of the slope coefficient estimates for each of the 10 countries, when it is used as an anchor country against the remaining 9 countries. There are several interesting findings in this table. First, as can be seen from the first row, the average slope coefficient estimates differ markedly from 1. In line with other empirical tests of UIP, they are negative for each of the 10 countries. Moreover, the standard deviation is larger than the average values for each country, implying that the estimates are very variable. In fact, the extreme values go from a minimum

[28]The number of possible combinations/country pairs in the set of 10 countries is 10x9x6 = 540. If we subtract the combinations/country pairs involving the 5 Eurozone countries in the last time subperiod (5x4=20), we obtain 540-20=520 combinations at the end.

of -16.03 to a maximum of 8.18. Finally, the figures in the lower panel point out that large differences exist even for the same anchor country across different time subperiods, thereby further strengthening the evidence of excessive variability in the slope coefficient estimates.

A graphical representation of the variability in the slope coefficient estimates is given in Figures 3.2, 3.3 and 3.4. Each of the nine lines corresponds to a different anchor country used in the conventional tests for the title (domestic) country. From these graphs it can be inferred that the slope coefficient estimates for the same (domestic) country differ markedly depending on the time sub-period and anchor (foreign) country used in the estimations. Moreover, large differences exist even among slope coefficient estimates for the same domestic country in the same time sub-period. In fact, they are typically higher (for most of the countries) when Italy or France is used as anchor country, and typically lower when the anchor country is Switzerland, the United States, Japan or Germany. This implies that besides variability among time sub-periods, slope coefficient estimates display large and systematic variability with respect to the anchor country included in the estimation. Therefore, our finding suggests that the results from these conventional tests might be related to the bias identified before.

Table 3.10: Summary Statistics: Slope Coefficient Estimates

Anchor	**BEL**	**CAN**	**FR**	**GER**	**ITA**	**JAP**	**NL**	**SWI**	**UK**	**USA**
mean	-1.10	-2.99	-0.39	-2.06	-0.05	-2.71	-1.92	-3.62	-1.88	-3.40
st.dev.	3.52	4.22	2.64	4.03	3.39	4.38	3.96	3.89	3.46	4.87
min	-13.05	-14.23	-7.77	-14.23	-7.29	-14.85	-16.03	-14.85	-13.58	-16.03
max	8.18	3.98	6.77	6.89	7.45	5.96	7.45	4.89	8.18	6.09

Anchor	**BEL**	**CAN**	**FR**	**GER**	**ITA**	**JAP**	**NL**	**SWI**	**UK**	**USA**
1975-1979	1.18	-2.27	0.10	-0.46	0.58	-0.74	-0.98	-2.76	0.11	0.05
1980-1984	-0.05	-1.36	0.38	-1.36	-0.57	-2.54	-1.52	-1.84	-3.48	-1.41
1985-1989	-1.75	-9.39	0.47	-5.27	-0.90	-5.35	-4.69	-4.86	-5.23	-8.61
1990-1994	-0.20	-1.19	0.72	0.13	5.02	-0.35	0.15	-1.95	-0.90	1.41
1995-1999	-3.39	0.87	-1.67	-2.98	-2.04	-6.57	-2.08	-4.44	-1.34	-6.08
2000-2004	-2.40	-4.64	-2.40	-2.40	-2.40	-0.72	-2.40	-5.90	-0.47	-5.83
1975-2004	-1.10	-2.99	-0.39	-2.06	-0.05	-2.71	-1.92	-3.62	-1.88	-3.40

Note:The table presents summary statistics for the slope coefficient estimates obtained from difference regressions, i.e. conventional empirical tests for the UIP condition. Each column in this table contains results from regressions that use the country indicated at the top of the column as anchor in the estimations (the name of the anchor country is mentioned at the top of the column). The upper panel contains summary statistics for the entire time period (January 1975-December 2004): average values for the slope coefficient estimates (first row), their standard deviations (second row), and extreme values (third and fourth row). The lower panel reports average values for the slope coefficient estimate in each of the six subperiods. Finally, the last row contains average values for the slope coefficient estimates over the entire period.

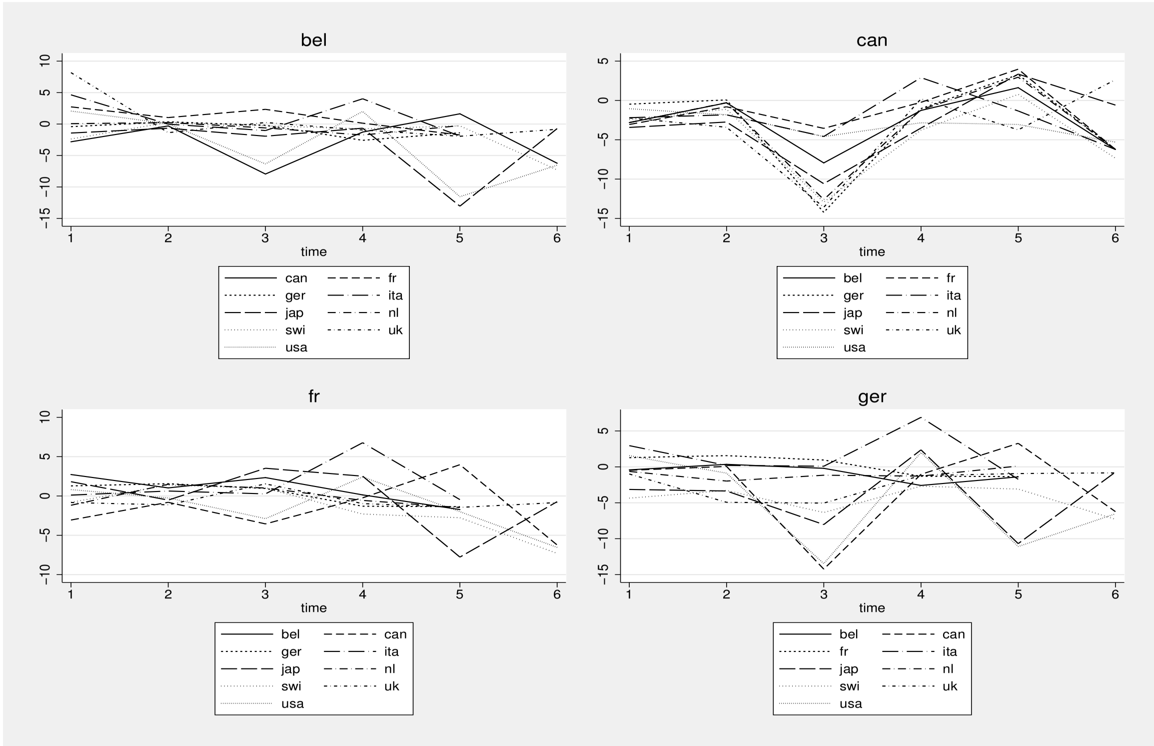


Figure 3.2: Estimates of the Slope Coefficient Using Different Anchor Countries

Note: The figure presents slope coefficient estimates obtained from difference regressions, i.e. conventional empirical tests for the UIP condition when alternative countries are used as anchors. Each graph contains slope coefficient estimates for one (domestic) country and each of the 9 lines corresponds to a different anchor country included in the estimations for that particular (domestic) country. The domestic country is indicated in the title of each graph, while the anchor country corresponding to each line is indicated in the legend. Time periods (1-6) are indicated on the horizontal axis (1975-1979(1), 1980-1984(2), 1985-1989(3), 1990-1994(4), 1995-1999(5) and 2000-2004(6)).

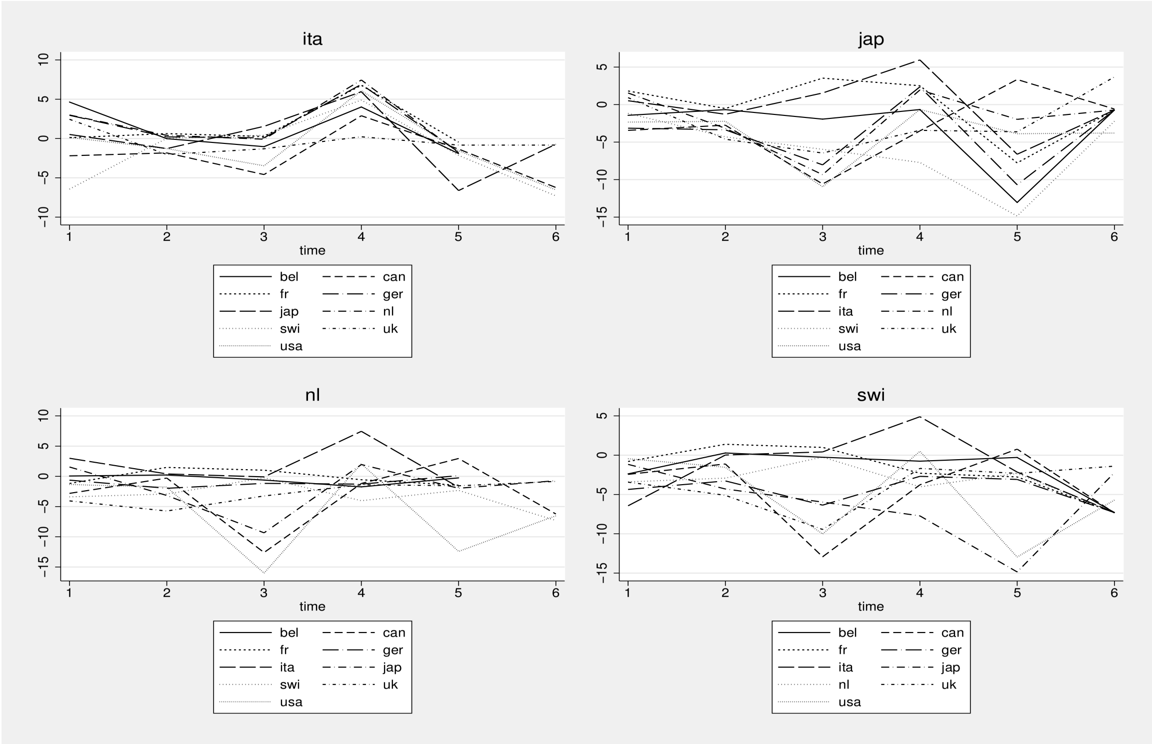


Figure 3.3: Estimates of the Slope Coefficient Using Different Anchor Countries

Note: The figure presents slope coefficient estimates obtained from difference regressions, i.e. conventional empirical tests for the UIP condition when alternative countries are used as anchors. Each graph contains slope coefficient estimates for one (domestic) country and each of the 9 lines corresponds to a different anchor country included in the estimations for that particular (domestic) country. The domestic country is indicated in the title of each graph, while the anchor country corresponding to each line is indicated in the legend. Time periods (1-6) are indicated on the horizontal axis (1975-1979(1), 1980-1984(2), 1985-1989(3), 1990-1994(4), 1995-1999(5) and 2000-2004(6)).

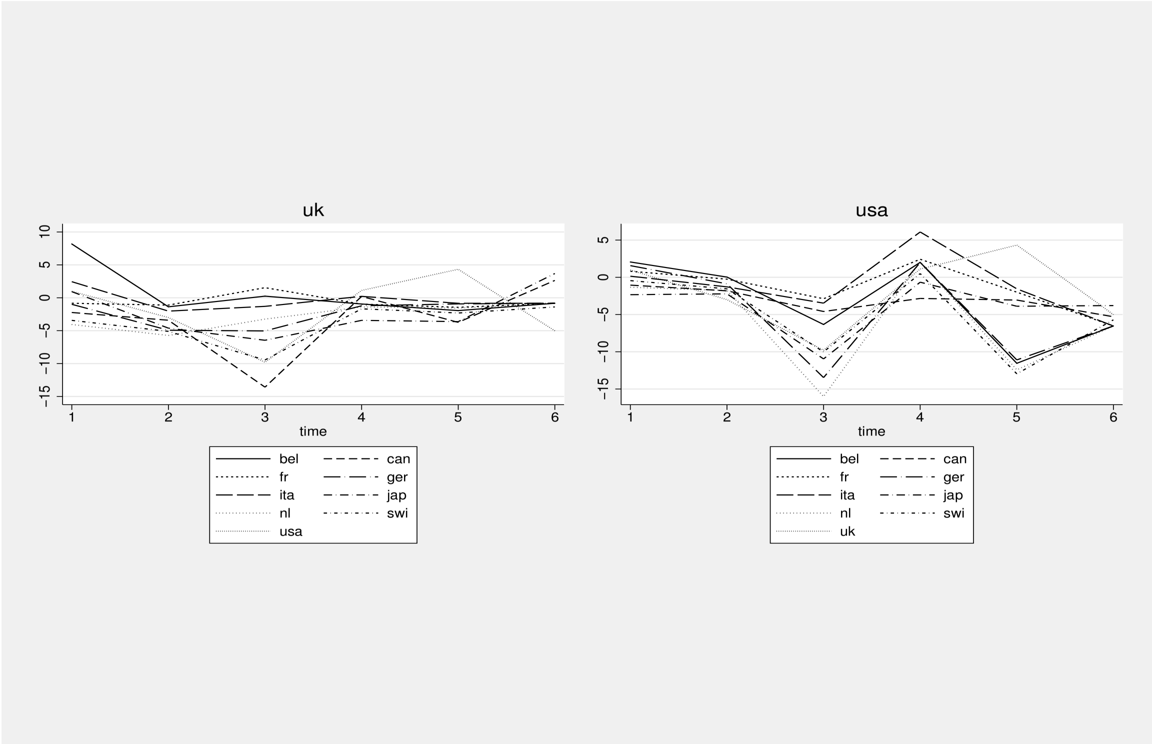


Figure 3.4: Estimates of the Slope Coefficient Using Different Anchor Countries

Note: The figure presents slope coefficient estimates obtained from difference regressions, i.e. conventional empirical tests for the UIP condition when alternative countries are used as anchors. Each graph contains slope coefficient estimates for one (domestic) country and each of the 9 lines corresponds to a different anchor country included in the estimations for that particular (domestic) country. The domestic country is indicated in the title of each graph, while the anchor country corresponding to each line is indicated in the legend. Time periods (1-6) are indicated on the horizontal axis (1975-1979(1), 1980-1984(2), 1985-1989(3), 1990-1994(4), 1995-1999(5) and 2000-2004(6)).

3.5 Can the Biasedness Hypothesis Account for the Anomalous Values of $\hat{\beta}$?

3.5.1 Symmetric Relationships among $\hat{\beta}$, δ, and Bias

As argued in section 3.2.4, the regression specification in equation (3.4) can only be an approximation of the true data-generating process. We chose it as it is the most simple and parsimonious way to demonstrate the potential effect of the omitted variable bias in the conventional UIP tests. However, due to the regression specification, the same slope coefficient estimate $\hat{\beta}_{ij} = \hat{\beta}_{ji}$ is obtained in the estimations for the country-pair (i,j) for two different δ's: the first one when country i is used as an anchor ($\delta = \frac{Var(i)}{Var(j)}$ in that case), and the second one when country j is used as anchor (therefore, $\delta = \frac{Var(j)}{Var(i)}$). This explains why the (upper) scatterplot in Figure 3.5 that depicts the delta-beta relationship is (quasi)symmetrical with respect to $\delta = 1$.

A similar argument explains the symmetrical pattern in the lower scatterplot in Figure 3.5. Since the total bias given by equation (3.9) is a function of δ, the argument above implies that the same slope coefficient estimate ($\hat{\beta}_{ij} = \hat{\beta}_{ji}$) will be associated with two total bias values.[29] In this scatterplot, same values for the slope coefficient estimate seem to appear on both sides of a point where total bias is negative, but close to 0. In fact, the scatterplot is indeed (quasi)symmetric with respect to the average value for total bias (-0.38).[30]

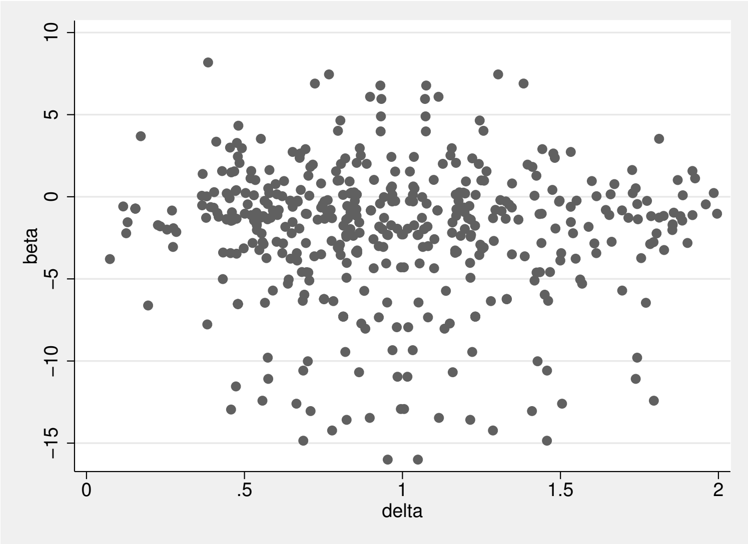
It follows straightforwardly that - due to the symmetry imposed by the empirical specification - a direct test of the theoretical relation between the bias value and $\hat{\beta}$ will yield insignificant results. In fact, regressing $\hat{\beta}_{ij}$ on the bias for all i and j results in an estimated slope coefficient of -0.0071 with t-statistic equal to -0.02.

However, dividing the total sample in two parts: one for all observations where $\delta < 1$, and one for those observations where $\delta > 1$ does lead to a strong effect of the bias on $\hat{\beta}$.[31] The slope coefficient for the first case ($\delta <$

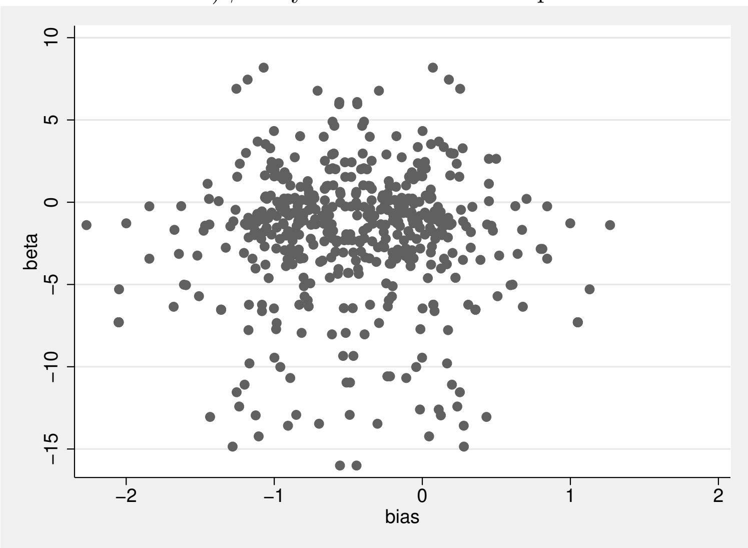
[29] One of these total bias expressions must be negative (when $\delta > 1$, see discussion in section 3.2.4), while the other might be positive.

[30] This is the average value for the total bias expression under the restriction that $\delta < 2$ (excludes 18 observations). The average bias for the entire sample is somewhat lower (-0.46).

[31] This separation of the dataset is similar to the separation with respect to $bias = 0$.



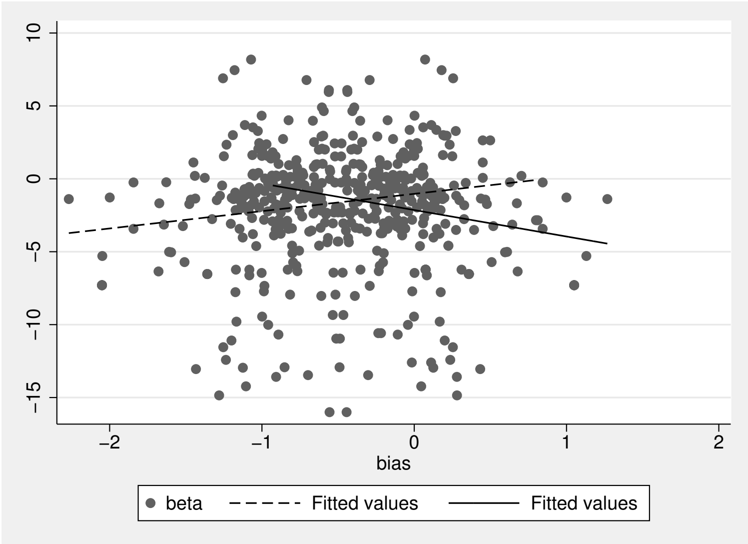
a) $\hat{\beta}$-δ Symmetric Relationship



b) $\hat{\beta}$-Bias Symmetric Relationship

Figure 3.5: Symmetric Relationships: $\hat{\beta}$-δ and $\hat{\beta}$-Bias

Note: The figure contains two scatterplots that depict the symmetric relationship between the slope coefficient estimates $\hat{\beta}$ and δ (upper panel) and the symmetric relationship between slope coefficient estimates $\hat{\beta}$ and total bias (lower panel).



Bias-$\hat{\beta}$ Relationship for $\delta > 1$ (dashed) and $\delta < 1$ (solid)

Figure 3.6: Total Bias and Slope Coefficient Estimates

Note: The figure contains a scatterplot that depicts the relationship between total bias (horizontal axis) and slope coefficient estimates $\hat{\beta}$ for all bilateral country-pairs included in the study (vertical axis). It includes two fitted lines: the solid line corresponds to the observations where $\delta < 1$, the dashed line corresponds to the observations where $\delta > 1$.

1) is -1.81 and is significant at the 1 percent significance level (t-statistic equals -2.59), while the slope coefficient for the second case ($\delta > 1$) is 1.18 and is significant at the 5 percent significance level (t-statistic equals 2.06). Figure 3.6 depicts the bias-$\hat{\beta}$ relationship for both parts of the sample. Hence, although no relationship can be found for all observations together due to the symmetric structure of the dataset, each of the two subsamples around the axis $\delta = 1$ suggests a strong effect of total bias on the slope coefficient estimates. Unfortunately, no inferences can be drawn from the sign of the coefficients due to the empirical setup.

3.5.2 The Importance of Interest Rate Volatility and ρ

The previous section demonstrated that the symmetric structure of our dataset does not allow for a direct test of the relationship between the bias term (or the relative volatility term δ) and the slope coefficient estimates. In order to avoid this symmetry, we focus on two alternative variables that were shown to be important for the bias term (section 3.2.4 and figure 3.1). Equation (3.9) in section 3.2.4 indicates that anchor (foreign) country interest rate volatility and the correlation between the interest rates series should theoretically have important effects on the bias term, and therefore, on the magnitude of the slope coefficient estimates.

In this section, we investigate to what extent this bias is due to the anchor country interest rate volatility and the coefficient of correlation between the interest rate series as suggested by the formal demonstration in section 3.2.4. For this purpose, we use two methods: visual inspection and formal (regression) analysis.[32]

Figure 3.7 displays the relationship between the average interest rate volatility for each country over the entire time period and the average slope coefficient estimates from regressions in which that country is used as an anchor. It shows a strongly positive relationship, suggesting that on average anchor countries with relatively more variable interest rates are associated with relatively higher slope coefficient estimates. For example, the anchor countries with the highest average interest rate volatility (Italy and France) are associated with the highest (average) values for the slope coefficient estimates. On the other hand, anchor countries with very low interest rate volatility (like Japan and Switzerland, for example) are associated with below-average values for the slope coefficient estimates. Obviously, anchor country interest rate volatility cannot explain the full amount of the bias in $\hat{\beta}$ as all (average) estimates displayed in Figure 3.7 are negative. Nevertheless, coefficients get significantly less anomalous (smaller in absolute value terms) for estimations with relatively higher anchor country interest rate volatility.[33]

We now turn to a formal analysis. In order to test for the effect of interest rate volatility on slope coefficient estimates, we estimate the following type of

[32]We only document the importance of anchor country interest rate volatility in the visual analysis part. The visual evidence for the coefficient of correlation is similar (opposite, negative relationships). Due to space reasons, we do not report these figures.

[33]Scatterplots for individual countries and for all observations (country-pairs) show a similar, though sometimes weaker pattern. They are available upon request.

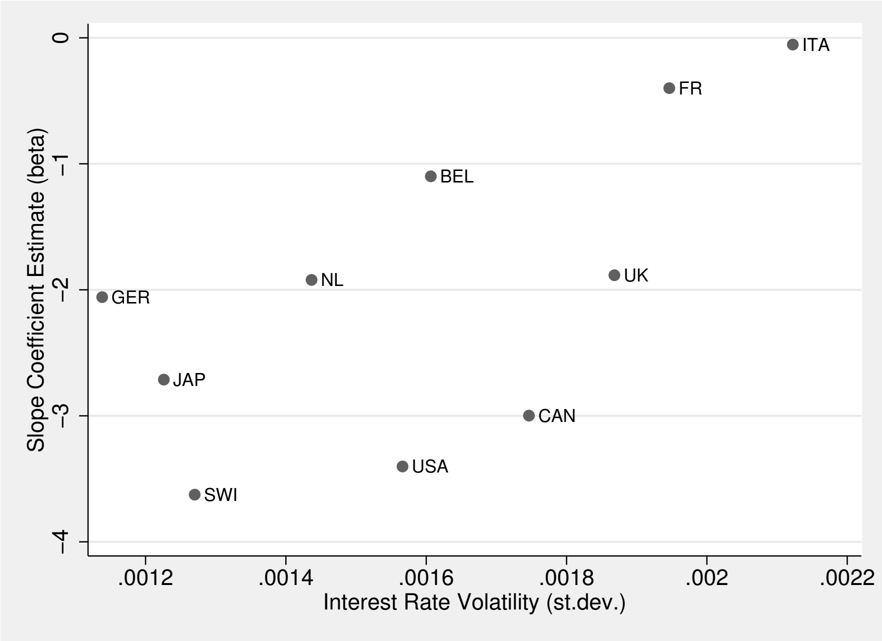


Figure 3.7: (Average) Interest Rate Volatility vs. (Average) Slope Coefficient Estimates

Note: The scatterplot in this figure displays the relationship between the average interest rate volatility for each country over the entire time period and the average slope coefficient estimates from regressions in which that country is used as an anchor. The sample period for Belgium, Italy, and Japan is July 1978-December 2004, while for all other countries is January 1975-December 2004. Average interest rate volatility refers to the average value for the interest rate standard deviation. It is calculated as the average value of the monthly standard deviations for each of the 6 time subperiods.

regression for each of the 10 countries i across subperiods t:

$$\hat{\beta}^{i}_{t,i^*} = \gamma_0 + \gamma_1 \sigma_{t,i^*} + \nu^{i}_{t,i^*} \tag{3.13}$$

where $\hat{\beta}^{i}_{t,i^*}$ is the slope coefficient estimate from the conventional UIP test for countries i (domestic) and i^* (anchor) for time sub-period t, σ_{t,i^*} is the interest rate standard deviation for the corresponding anchor country i^* for time sub-period t, and $\nu^{i}_{t,i^*} \sim iid(0, \sigma^2_{\nu^i})$ is the error term.

Similarly, in order to test for the effect of the coefficient of correlation between the two (domestic and anchor country) interest rate series on the slope coefficient estimates, we estimate the following type of regressions:

$$\hat{\beta}^i_{t,i^*} = \phi_0 + \phi_1 \rho_t + \eta^i_{t,i^*} \tag{3.14}$$

where ρ_t is the coefficient of correlation at time t, and $\eta^i_{t,i^*} \sim iid(0, \sigma^2_{\nu^i})$ is the error term. The results from the regressions given by equation (3.13) are displayed in the upper panel of Table 3.11. The slope coefficient estimate γ_1 is positive for 9 out of 10 countries (the only exception being the regression for UK). Moreover, it is significant at the 1 percent significance level for 2 countries (Switzerland and USA), at 5 percent significance level for 2 other countries (Belgium and Germany), and marginally (in)significant for Japan and the Netherlands. The last column of Table 3.11 reports the results from the pooled-OLS regression for all 10 countries. The slope coefficient is positive and significant at any conventional significance level (t-statistic equals 5.22). These regression results confirm the visual evidence shown in the previous section, suggesting that a (strong) positive relationship exists between interest rate volatility and slope coefficient estimates from conventional tests of UIP.[34]

The lower panel of Table 3.11 presents results for the regressions given by equation (3.14). The slope coefficient ϕ_1 is negative for each country. Moreover, for four countries it is significant at least at the 5 percent, and for the USA even at the 1 percent significance level.[35] Finally, the last column shows results for the complete sample: ϕ_1 is significant even at the 1 percent level in this case (t-statistic equals -4.83). In sum, the evidence suggests a strongly negative relationship between ρ and $\hat{\beta}$: very low (negative) values for $\hat{\beta}$ are typically associated with high values for ρ.[36] Overall, the evidence suggests that $\hat{\beta}$ gets increasingly negative when ρ increases and σ_{i^*} decreases. This is consistent with the theoretical hypothesis in section 3.2.4 and Figure 3.1. Put differently, the more interest rates behave alike in correlation and volatility, the larger the (downward) bias on $\hat{\beta}$.

[34] The results are robust to the inclusion of country-pair-specific fixed-effects, exclusion of certain time subperiods or anchor countries. All results are available upon request.

[35] These same four countries (Belgium, Germany, Switzerland, and the USA) had significant results in the upper panel as well.

[36] Note that these results are broadly in line with Haynes and Stone (1981, 1982), who demonstrate that the specification bias introduced in difference regressions is more likely to lead to slope coefficient sign reversal(s) when the underlying variables are strongly (positively) correlated.

Table 3.11: Effects of Anchor Country Interest Rate Volatility and Interest Rate Correlation on Slope Coefficient Estimates

	BEL	CAN	FR	GER	ITA	JAP	NL	SWI	UK	USA	ALL
σ_{t,i^*}	**1,216.17**	503.59	558.05	**1,111.15**	689.04	962.87	834.01	**1,533.59**	-256.77	**1,954.46**	**895.08**
	(2.52)**	(0.88)	(1.46)	**(2.11)****	(1.41)	(1.63)	(1.57)	**(3.22)*****	(-0.53)	**(3.20)*****	**(5.22)*****
Constant	-3.034	-3.792	-1.266	-3.884	-1.111	-4.285	-3.264	-6.123	-1.483	-6.52	-3.44
	(-3.39)***	(-3.55)***	(-1.83)*	(-3.83)***	(-1.26)	(-3.80)***	(-3.24)***	(-6.67)***	(-1.67)	(-5.67)***	(-10.71)***
Observations	50	54	50	50	50	54	50	54	54	54	520
R-squared	0.11	0.01	0.04	0.08	0.04	0.05	0.05	0.17	0.01	0.16	0.048
ρ	**-3.881**	-1.020	-1.444	**-5.284**	-0.658	-0.802	-1.339	**-3.531**	-1.866	**-8.234**	**-3.169**
	(-2.05)**	(-0.43)	(-1.03)	**(-2.13)****	(-0.31)	(-0.36)	(-0.56)	**(-2.26)****	(-1.01)	**(-3.56)*****	**(-4.83)*****
Constant	1.112	-2.379	0.312	1.279	0.270	-2.241	-1.209	-1.524	-0.843	1.232	-0.247
	(0.91)	(-1.54)	(0.38)	(0.73)	(0.23)	(-1.57)	(-0.73)	(-1.44)	(-0.74)	(0.86)	(-0.59)
Observations	50	54	50	50	50	54	50	54	54	54	520
R-squared	0.08	0.01	0.02	0.09	0.01	0.01	0.01	0.09	0.02	0.20	0.04

Note: The upper panel of this table presents pooled-OLS estimation results from the following regression equations: $\hat{\beta}^i_{t,i^*} = \gamma_0 + \gamma_1 \sigma_{t,i^*} + \nu^i_{t,i^*}$, where $\hat{\beta}^i_{t,i^*}$ is the slope coefficient estimate from a conventional (bilateral) UIP test between country i and anchor country i^* in time period t, and σ_{t,i^*} is the standard deviation of the interest rate for anchor country i^* in time period t. The lower panel of this table presents pooled-OLS estimation results from the following regression equations: $\hat{\beta}^i_{t,i^*} = \phi_0 + \phi_1 \rho_t + \eta^i_{t,i^*}$, where ρ_t is the coefficient of (simple) correlation between the two interest rate series i^* and i^* at time t. The last column contains estimation results for the entire dataset (all observations). t-statistics are reported in parentheses below the slope coefficient estimates. Significance at 10%, 5%, and 1% level is indicated by *, **, and ***, respectively. Bilateral regressions results between countries from the Eurozone in the last time subperiod 2000-2004 are excluded from the analysis. Hence, the estimation results for these countries are based on 50 observations (not 54 as for the other countries).

3.6 Concluding Remarks

Our analysis suggests that the slope coefficient estimator in conventional empirical tests of UIP may be biased if the true data generating process is not exactly equal to the theoretically expected one. Using several alternative specifications, we demonstrate that the rejection of the UIP might be limited to the conventional empirical tests that regress the change in the nominal exchange rate on the interest rate differential. First, the results from level regressions suggest that, when expressed in common currency terms, the holding period returns of deposits denominated in different currencies move very closely together. In fact, the slope coefficients in these specifications do not significantly differ from unity. Second, the (anchor country) interest rate enters significantly the conventional specification of UIP, suggesting that its omission will generally lead to biased estimation results. Furthermore, the implied values for the "true β" from these regressions or the slope coefficient estimates from the corresponding constrained regressions generally fall in the range 0.4-0.9. We interpret these findings as suggestive evidence that although the "true" slope coefficient might be close to unity, the empirical procedure used in conventional tests of UIP might be biased and unable to account for it.

We demonstrate that the symmetric structure of our dataset does not allow for a direct test of the relationship between the bias and the slope coefficient estimates. Instead, we separately test for this relationship on each side of the symmetry-axis and find significant results on both sides. Moreover, in order to avoid this symmetry, we focus on two determinants for the bias term: the anchor (foreign) country interest rate volatility and the correlation between the interest rates series. In this way, we (indirectly) test for the importance of the bias (determinants) for the slope coefficient estimates. The empirical evidence provides strong support for our argument: first, higher (anchor country) interest rate volatility is associated with higher slope coefficient estimates 9 out of 10 countries); and second, higher interest rate series correlation is strongly associated with lower slope coefficient estimates (all 10 countries).

A note of caution is worth for the interpretation of the results from this study. Although we find some evidence about the importance of the bias (determinants) on the slope coefficient estimates, we do not manage to "recover" values required by the UIP. Actually, the average slope coefficient estimates stay negative for each country in the analysis. We conclude that these findings correspond with the main goal of the chapter - not to attempt to "rehabilitate" the UIP nor solve the UIP puzzle, but rather to provide evidence about the estimation bias related to the conventional UIP testing procedure.

Several recent empirical studies attempt to "rehabilitate" UIP using different strategies. Therefore, it might be interesting to relate the estimation bias with the support for UIP in tests using longer horizons (Chinn and Meredith, 2005; Chinn, 2006), focusing on the very short maturity spectrum (Chaboud and Wright, 2005), or including developing economies (Bansal and Dahlquist, 2000; Frankel and Poonawala, 2004; Chinn, 2006). Moreover, the empirical evidence against UIP is mainly based on studies that are centered on the US (dollar) and include other G-7 (or OECD) countries only. Therefore, the very similar interest rate series for all these countries might be one possible explanation for the anomalous findings. In this way, it might be interesting to relate the results from this study with the work by Lustig and Verdelhan (2007), where UIP deviations are explained by currency (country) portfolios based on their interest rate differences against the US dollar.

Appendix A

A.1 Stationarity Tests

The complete results from the stationarity tests are reported in Tables A.1, A.2, A.3, and A.4. Each table refers to one variable used in the difference regressions (Tables A.1 and A.2)or the level regressions (Tables A.3 and A.4).[1]

Table A.1 contains tests statistics from the unit-root tests on the change in the nominal exchange rate, which is used as dependent variable in the difference (conventional) regressions.[2] The results from the ADF tests in this table suggest that the null hypothesis of unit-root (nonstationarity) in the nominal exchange rate changes can be rejected at any conventional significance level. The null hypothesis in the general ADF tests (no drift, no trend) is always rejected at the 10 percent significance level. Moreover, if we allow for drift parameter in the test specification, then (almost) all test statistics suggest rejection even at the 1 percent significance level.[3] The last column in this table reports test statistics from the panel unit-root tests following Maddala and Wu (1999). All statistics in this column suggest very strong rejection of the null hypothesis of nonstationarity even at 0.1 percent significance level.[4]

[1]The series tested here correspond exactly with the data used in the estimations presented in tables 3.1-3.4.

[2]The maximum number of lags in these tables (lags= 16) is selected according to the Schwert (1989) criterion.

[3]The inclusion of a drift parameter in this testing procedure corresponds with the allowance of a non-zero expected value for the change in the nominal exchange rate.

[4]The Maddala-Wu test statistics reported in this column are calculated from the simple ADF tests (no drift, no trend). The corresponding statistics calculated using the results from the ADF tests with drift (not reported in the table) suggest even stronger rejection of the nonstationarity null hypothesis.

The results for the interest rate difference that is used as independent variable in conventional UIP tests are reported in Table A.2. Generally, the test statistics from the simple ADF time-series tests suggest rejection of the nonstationarity hypothesis at the 10 percent significance.[5] Furthermore, including a drift parameter in the basic ADF specification improves the results significantly: all test statistics are significant at least at the 10 percent, and most of them at the 1 percent significance.[6] A similar picture is seen in the last two columns of this table: the panel unit-root hypothesis without drift can only be rejected at low lag orders (5 lags at the most). However, inclusion of drift parameter improves the significance of the results substantially (last column): the panel unit-root is rejected at any reasonable significance level. In sum, both variables used in the conventional (difference) regressions for testing the empirical validity of UIP seem to be stationary. Now, we turn to the stationarity properties of the variables used in the "unconventional" or level regressions reported in tables 3.3 and 3.4.

Table A.3 contains results from the unit-root tests for the dependent variable in the level regressions - the domestic currency holding period returns of US dollar deposits. The results are very similar to those for the change in the nominal exchange rate reported in Table A.1, implying that the exchange rate changes constitute the core part of (the variability in) the domestic currency holding period returns of US dollar deposits. The simple ADF (no drift, no trend) results indicate rejection of the null hypothesis in almost all cases at the 10 percent significance level. Moreover, including a drift parameter in the specification leads to a substantial increase in statistical significance - all test statistics suggest rejection of the null hypothesis even at the 1 percent significance level. Furthermore, the results from the panel unit-root tests reported in the last column just strengthen the evidence against the presence of a unit-

[5] Most of the cases where nonstationarity cannot be rejected concern the following four countries: Germany, Italy, the Netherlands, and Switzerland.

[6] Inclusion of a drift parameter in these tests might be justified on several grounds. For example, the interest rate differential for two countries usually has a non-zero mean (see summary statistics) even over extended periods of time. Hence, accounting for this pattern (positive or negative expected interested rate differential) in the series might explain why the test statistics for Germany, Italy, the Netherlands, and Switzerland gain significance in the latter case.

root in this series. Similar as in the test results for the nominal exchange rate changes (Table A.1), the panel unit-root null hypothesis is rejected even at 0.1 percent significance level.[7] All these results imply normal, stationary behavior of the dependent variable in the level (unconventional) regressions.

Table A.4 contains results from the unit-root tests of the (domestic) interest rate series that are used as independent variables in the level regressions. The tests statistics from the ADF time-series tests offer mixed evidence about the stationarity of the series. When no drift is included in the testing specification, the null hypothesis only is rejected in a very limited number of cases.[8] However, the introduction of a drift parameter in the test equation strengthens the evidence against the null hypothesis of nonstationarity: it is rejected for most countries (at most lag orders) at the 5 or 10 percent significance level. Nonetheless, for certain countries it still cannot be rejected at any conventional significance level. For example, in the case of Italy the null hypothesis of unit-root cannot be rejected even at the 10 percent significance for most lag orders (when more than 4 lags are included).

In order to test for nonstationarity in the panel, we present the results from the panel unit-root tests in the last four columns in Table A.4. The first two of these columns refer to tests based on specifications without drift, while the last two columns refer to tests that allow for drift parameter in the baseline specification. Moreover, for each of these specifications, we report results from unit-root tests based on two different datasets: the first one includes all countries/observations (complete panel), while the second one excludes Italy. The time-series unit-root tests for Italy implied non-rejection of the unit-root null hypothesis at any conventional significance level. Therefore, by excluding Italy from the second dataset, we eliminate the possibility that the test results are driven by an “outlier”, and test for presence of a panel unit-root in the interest rate series for the other 8 countries in the dataset.

The results based on specifications without drift parameter, reported in

[7] Again, the Madda-Wu statistics presented in this table are calculated using the p-values from the simple ADF unit-root tests (no trend, no drift). The results calculated from ADF unit-root tests with drift are even less favorable for the null hypothesis at hand.

[8] Moreover, the cases of rejection generally refer to two countries only: France and the Netherlands. For most countries, the null hypothesis cannot be rejected at any lag order and any reasonable significance level.

the first two columns, imply that in general the null hypothesis cannot be rejected even at the 10 percent significance level. On the other hand, the results in the last two columns are based on tests that allow for drift term in the specification, and generally lead to rejection of the null hypothesis that (domestic) interest rates are nonstationary. The latter point is true for the panel unit-root tests based on the entire dataset as well as those that exclude Italy. In the latter case (when Italy is excluded for the dataset), the evidence against nonstationarity improves substantially: first, the null hypothesis is rejected at all but one lag order (16 lags), and second, the results gain in significance, so that the null hypothesis is rejected at the 5 percent level for most lag orders.

In sum, there is not much evidence that the time-series used in level and/or difference regressions presented in tables 3.3-3.2 follow nonstationary processes. In fact, most results from the unit-root tests on the variables used in the difference regressions (presented in tables A.1 and A.2) confirm the finding that conventional UIP tests are based on stationary variables. Furthermore, the results from the unit-root tests on the variables used in the level (unconventional) regressions (presented in tables A.3 and A.4) do not give much evidence against stationarity. In fact, the only mixed evidence is found for the (domestic) interest rate variable. However, the evidence against stationarity seems to be limited to individual countries (Italy). When we include all countries and test for panel unit-roots, the null hypothesis is generally rejected at the 10 percent level. Finally, excluding one nonstationary time-series from the panel leads to more significant results, and in turn, stronger evidence that interest rates might follow stationary process with drift.[9] In sum, we conclude that nonstationarity is not a (very) serious issue for the variables concerned, and therefore, most of the estimation results from the level regressions (tables 3.3 and 3.4) can be considered valid.

[9] In order to see whether the panel data results are "driven" by the nonstationary interest rate for Italy, we estimated panel model excluding Italy. As shown in the last two columns of Table 3.4, the results from the panel estimations do not change a lot.

Table A.1: Unit-Root Test Results for the Change in the Nominal Exchange Rate

country	BEL	CAN	FRA	GER	ITA	JAP	NL	SWI	UK	ALL (Panel)
lag(s)	ADF	ADF	ADF	ADF	ADF	ADF	ADF	ADF	ADF	MW
0	-18.935***	-19.688***	-18.96***	-18.999***	-17.702***	-19.042***	-18.762***	-18.759***	-17.992***	90.000****
1	-12.223***	-14.29***	-12.477***	-12.391***	-11.987***	-12.781***	-12.134***	-12.299***	-12.884***	90.000****
2	-10.034***	-11.287***	-9.949***	-10.244***	-9.83***	-10.286***	-10.175***	-10.484***	-10.951***	90.000****
3	-9.338***	-9.577***	-9.146***	-9.536***	-9.235***	-8.559***	-9.578***	-9.773***	-9.386***	90.000****
4	-8.244***	-8.282***	-8.117***	-8.391***	-7.979***	-8.437***	-8.386***	-8.475***	-8.213***	90.000****
5	-7.73***	-7.429***	-7.581***	-7.879***	-7.518***	-8.462***	-7.824***	-8.21***	-7.914***	90.000****
6	-6.569***	-7.281***	-6.456***	-6.779***	-6.612***	-7.741***	-6.813***	-7.069***	-7.408***	90.000****
7	-5.852***	-6.529***	-5.863***	-6.155***	-5.895***	-6.533***	-6.172***	-6.625***	-6.632***	90.000****
8	-5.115***	-5.66***	-5.03***	-5.394***	-5.248***	-5.562***	-5.389***	-5.62***	-5.973***	90.000****
9	-4.876***	-4.765***	-4.996***	-5.169***	-4.872***	-5.574***	-5.228***	-5.667***	-6.022***	88.000****
10	-4.191***	-3.938***	-4.333***	-4.418***	-4.306***	-4.767***	-4.448***	-4.771***	-5.036***	65.835****
11	-4.184***	-3.63***	-4.234***	-4.322***	-4.304***	-4.493***	-4.382***	-4.683***	-4.924***	63.357****
12	-4.012***	-3.789***	-4.064***	-4.171***	-4.253***	-4.671***	-4.236***	-4.793***	-4.85***	62.091****
13	-3.889***	-3.718***	-4.007***	-4.06***	-4.265***	-4.517***	-4.134***	-4.548***	-5.136***	59.390****
14	-4.005***	-3.826***	-4.093***	-4.154***	-4.407***	-4.316***	-4.303***	-4.419***	-5.229***	60.755****
15	-3.553***	-3.227**	-3.676***	-3.812***	-4.185***	-4.268***	-3.872***	-4.061***	-5.036***	51.757****
16	-3.722***	-3.34**	-3.762***	-4.067***	-4.111***	-4.587***	-4.121***	-4.292***	-5.006***	56.265****

Note: The table reports Augmented Dickey-Fuller (ADF) and and Maddala-Wu (MW) test statistics based on the null hypothesis that the series follows a unit-root process over the period January 1975-December 2004. The nominal exchange rate is defined as the number of domestic currency units per US dollar, and all changes in the nominal exchange rate are calculated relative to the US dollar at monthly frequency. The maximum number of lags (lags= 16) is selected according to the Schwert (1989) criterion. The critical ADF (374 observations, without trend, without drift) values for rejection of the null hypothesis are given as follows: -2.57 (10% significance level), -2.875 (5%), -3.45 (1%). When allowing for drift, the values of the test statistics stay the same, but now they are compared against the critical values of the Student's t-distribution. The (lower) critical values for two-sided t-test with 358-374 degrees of freedom are given as follows: -1.284 (10% significance level), -1.649 (5%), -2.337 (1%). The last column reports panel unit-root test statistics calculated according to Maddala and Wu (1999). Maddala and Wu (1999) panel unit-root test is one sided test that follows χ^2-distribution with $2n$ degrees of freedom (where n is the number of separate time-series). The (upper) critical values for the χ^2-distribution with 18 degrees of freedom are given as follows: 25.989 (10% significance level), 28.869 (5%), 34.805 (1%), 42.312 (0.1%).
Bold faces refer to cases where the null hypothesis of nonstationarity (presence of unit root) of the series without allowing for drift is rejected at least at the 10% significance level. When a drift is included in the tests, rejection of the null hypothesis at 10%, 5% and 1% significance level is indicated by *, **, and ***, respectively. In the case of Maddala-Wu panel unit-root tests reported in the last column, **bold** faces refer to cases where the null hypothesis of nonstationarity (presence of panel unit-root) is rejected at least at the 10% significance level. Moreover, significance at 10%, 5%, 1%, and 0.1% is indicated by *, **, ***, and ****, respectively.

Table A.2: Unit-Root Test Results for the Interest Rate Differential

country	BEL	CAN	FRA	GER	ITA	JAP	NL	SWI	UK	ALL (Panel)	ALL (Panel)
lag	ADF	ADF	ADF	ADF	ADF	ADF	ADF	ADF	ADF	MW (no drift)	MW (with drift)
0	**-5.996*****	**-4.841*****	**-7.259*****	**-2.693*****	**-5.525*****	**-3.224*****	**-4.046*****	**-2.727*****	**-4.187*****	**60.175*******	**76.136*******
1	**-4.310*****	**-4.282*****	**-4.760*****	**-3.334*****	**-4.206*****	**-3.661*****	**-4.392*****	**-2.722*****	**-4.318*****	**52.392*******	**79.539*******
2	**-3.427*****	**-3.914*****	**-4.632*****	-2.426***	**-4.469*****	**-3.620*****	**-3.812*****	**-2.735*****	**-3.951*****	**44.087*******	**69.611*******
3	**-3.393*****	**-3.665*****	**-4.349*****	-2.354***	**-3.856*****	**-3.620*****	**-3.711*****	-2.527***	**-4.079*****	**39.485*****	**66.673*******
4	**-3.330*****	**-3.013*****	**-4.021*****	-2.005**	**-3.506*****	**-3.729*****	**-2.957*****	-2.029**	**-3.747*****	**31.471****	**57.510*******
5	**-3.097*****	**-2.833*****	**-3.167*****	-2.082**	**-2.809*****	**-3.815*****	**-2.879*****	-1.838**	**-3.943*****	**26.852***	**52.269*******
6	**-3.337*****	**-2.660*****	**-2.783*****	-1.565*	-2.370***	**-3.424*****	-2.499***	-1.713**	**-3.521*****	21.459	**44.545*******
7	**-3.246*****	-2.298**	**-2.750*****	-1.383*	-2.454***	**-2.594*****	-2.542***	-1.595*	**-3.601*****	18.779	**40.793*****
8	**-2.847*****	-2.441***	-2.393***	-1.654**	-2.119**	**-2.765*****	-2.415***	-1.820**	**-3.497*****	17.124	**38.967*****
9	**-2.920*****	**-3.134*****	**-2.875*****	-2.014**	-2.245**	**-3.293*****	**-2.837*****	-1.943**	**-3.838*****	23.547	**46.822*******
10	**-3.023*****	**-2.758*****	-2.449***	-1.930**	-1.672**	**-3.251*****	-2.546***	-1.968**	**-3.294*****	19.002	**41.497*****
11	**-3.024*****	**-3.005*****	-2.386***	-1.886**	-1.332*	**-2.697*****	-2.504***	-1.633*	**-3.474*****	17.856	**39.512*****
12	**-2.844*****	**-2.818*****	**-2.667*****	-1.713**	-1.528*	**-2.702*****	-2.142**	-1.518*	**-2.709*****	14.969	**36.058*****
13	**-2.972*****	**-2.588*****	-2.545***	-1.954**	-1.384*	-2.468***	-2.159**	-1.462*	**-2.663*****	14.158	**34.969****
14	**-3.278*****	**-2.994*****	**-2.818*****	-1.932**	-1.530*	**-2.843*****	-2.254**	-1.628*	**-2.862*****	17.850	**39.721*****
15	**-3.452*****	**-3.104*****	**-3.197*****	-2.192**	-1.923**	**-3.373*****	-2.218**	-1.610*	**-2.625**	20.966	**43.788*******
16	**-3.989*****	**-3.202*****	**-3.046*****	**-2.729*****	-1.515*	**-3.469*****	-2.174**	-1.740**	**-2.804**	23.644	**47.999*******

Note: The table reports Augmented Dickey-Fuller (ADF) and and Maddala-Wu (MW) test statistics based on the null hypothesis that the series follows a unit-root process. Due to limitations in the interest rate data, the tests for Belgium, Italy, and Japan are conducted over the period July 1978-December 2004. For all other countries, the tests are conducted over the entire period January 1975-December 2004. All interest rate differences are calculated relative to the US one-month Eurocurrency deposit interest rate. The maximum number of lags (lags= 16) is selected according to the Schwert (1989) criterion. The critical ADF (332 observations, without trend, without drift) values for rejection of the null hypothesis are given as follows: -2.57 (10% significance level), -2.875 (5%), -3.45 (1%). When allowing for drift, the values of the test statistics stay the same, but now they are compared against the critical values of the Student's t-distribution. The (lower) critical values for two-sided t-test with 316-332 degrees of freedom are given as follows: -1.284 (10% significance level), -1.649 (5%), -2.337 (1%). The last two columns report panel unit-root test statistics calculated according to Maddala and Wu (1999): the first column presents results for tests without drift, while the second column presents results for tests that allow for drift parameter. Maddala and Wu (1999) panel unit-root test is one sided test that follows χ^2-distribution with $2n$ degrees of freedom (where n is the number of separate time-series). The (upper) critical values for the χ^2-distribution with 18 degrees of freedom are given as follows: 25.989 (10% significance level), 28.869 (5%), 34.805 (1%), 42.312 (0.1%).

Bold faces refer to cases where the null hypothesis of nonstationarity (presence of unit root) of the series without allowing for drift is rejected at least at the 10% significance level. When a drift is included in the tests, rejection of the null hypothesis at 10%, 5% and 1% significance level is indicated by *, **, and ***, respectively. In the case of Maddala-Wu panel unit-root tests reported in the last two columns, **bold** faces refer to cases where the null hypothesis of nonstationarity (presence of panel unit-root) is rejected at least at the 10% significance level. Moreover, significance at 10%, 5%, 1%, and 0.1% is indicated by *, **, ***, and ****, respectively.

Table A.3: Unit-Root Test Results for Domestic Currency Holding Period Returns of US Dollar Deposits

country	BEL	CAN	FRA	GER	ITA	JAP	NL	SWI	UK	ALL (Panel)
lag	ADF	ADF	ADF	ADF	ADF	ADF	ADF	ADF	ADF	MW (no drift)
0	-18.336***	-18.376***	-18.319***	-18.489***	-17.136***	-18.685***	-18.285***	-18.384***	-17.512***	90.000****
1	-11.591***	-12.975***	-11.779***	-11.835***	-11.439***	-12.364***	-11.566***	-11.872***	-12.379***	90.000****
2	-9.454***	-10.020***	-9.349***	-9.733***	-9.298***	-9.909***	-9.652***	-10.096***	-10.423***	90.000****
3	-8.708***	-8.222***	-8.512***	-8.990***	-8.658***	-8.208***	-9.026***	-9.347***	-8.859***	90.000****
4	-7.601***	-6.966***	-7.427***	-7.836***	-7.398***	-8.063***	-7.814***	-8.071***	-7.740***	90.000****
5	-7.096***	-6.118***	-6.836***	-7.345***	-6.923***	-8.047***	-7.273***	-7.810***	-7.382***	90.000****
6	-5.971***	-5.877***	-5.855***	-6.253***	-6.063***	-7.321***	-6.266***	-6.634***	-6.853***	90.000****
7	-5.322***	-5.168***	-5.292***	-5.659***	-5.391***	-6.177***	-5.672***	-6.219***	-6.158***	90.000****
8	-4.576***	-4.411***	-4.526***	-4.909***	-4.765***	-5.260***	-4.900***	-5.225***	-5.431***	80.444****
9	-4.403***	-3.670***	-4.461***	-4.739***	-4.413***	-5.271***	-4.777***	-5.293***	-5.546***	72.183****
10	-3.76***	-3.000***	-3.857***	-4.018***	-3.886***	-4.504***	-4.035***	-4.406***	-4.582***	52.532****
11	-3.724***	-2.747***	-3.747***	-3.890***	-3.970***	-4.240***	-3.950***	-4.320***	-4.468***	49.233****
12	-3.568***	-2.900***	-3.583***	-3.762***	-3.804***	-4.404***	-3.818***	-4.401***	-4.389***	47.817****
13	-3.453***	-2.832***	-3.535***	-3.647***	-3.916***	-4.270***	-3.711***	-4.158***	-4.755***	46.553****
14	-3.555***	-2.890***	-3.580***	-3.743***	-4.011***	-4.098***	-3.851***	-4.034***	-4.687***	47.129****
15	-3.195***	-2.418***	-3.236***	-3.468***	-3.498***	-4.072***	-3.502***	-3.716***	-4.518***	39.245***
16	-3.325***	-2.505 ***	-3.298***	-3.666***	-3.581***	-4.375***	-3.686***	-3.929***	-4.295***	42.227***

Note: The table reports Augmented Dickey-Fuller (ADF) and and Maddala-Wu (MW) test statistics based on the null hypothesis that the series follows a unit-root process over the period January 1975-December 2004. The holding period returns are calculated as the sum of nominal exchange rate changes relative to the US dollar (monthly frequency) and the one-month Eurocurrency deposit rate on the US dollar. The nominal exchange rate is defined as the number of domestic currency units per US dollar. The maximum number of lags (lags= 16) is selected according to the Schwert (1989) criterion. The critical ADF (374 observations, without trend, without drift) values for rejection of the null hypothesis are given as follows: -2.57 (10% significance level), -2.875 (5%), -3.45 (1%). When allowing for drift, the values of the test statistics stay the same, but now they are compared against the critical values of the Student's t-distribution. The (lower) critical values for two-sided t-test with 358-374 degrees of freedom are given as follows: -1.284 (10% significance level), -1.649 (5%), -2.337 (1%). The last column reports panel unit-root test statistics calculated according to Maddala and Wu (1999). Maddala and Wu (1999) panel unit-root test is one sided test that follows χ^2-distribution with $2n$ degrees of freedom (where n is the number of separate time-series). The (upper) critical values for the χ^2-distribution with 18 degrees of freedom are given as follows: 25.989 (10% significance level), 28.869 (5%), 34.805 (1%), 42.312 (0.1%).

Bold faces refer to cases where the null hypothesis of nonstationarity (presence of unit root) of the series without allowing for drift is rejected at least at the 10% significance level. When a drift is included in the tests, rejection of the null hypothesis at 10%, 5% and 1% significance level is indicated by *, **, and ***, respectively. In the case of Maddala-Wu panel unit-root tests reported in the last column, **bold** faces refer to cases where the null hypothesis of nonstationarity (presence of panel unit-root) is rejected at least at the 10% significance level. Moreover, significance at 10%, 5%, 1%, and 0.1% is indicated by *, **, ***, and ****, respectively.

Table A.4: Unit-Root Test Results for Interest Rates (Returns) in Domestic Currency

country	BEL	CAN	FRA	GER	ITA	JAP	NL	SWI	UK	Panel	Panel(ex.Ita)	Panel	Panel(ex.Ita)
lag	ADF	ADF	ADF	ADF	ADF	ADF	ADF	ADF	ADF	MW (no drift)		MW (with drift)	
0	**-4.028*****	-1.552*	**-4.608*****	-1.526*	**-3.150*****	-1.640*	**-3.258*****	**-2.834*****	-2.299**	**26.485***	23.213	**49.097******	**43.005******
1	**-2.753*****	-1.720**	**-2.994*****	-1.396*	-2.120**	-1.211	**-3.496*****	-1.822**	-2.100**	14.361	13.108	**34.425****	**30.906****
2	-1.99**	-1.915**	**-2.902*****	-1.284*	-2.122**	-1.209	**-3.193*****	-1.833**	-1.908**	11.959	10.704	**31.385****	**27.861****
3	-1.874**	-1.887**	**-2.646*****	-1.556*	-1.973**	-1.385*	**-3.490*****	-2.230**	-2.130**	12.969	11.919	**32.983****	**29.768****
4	-1.999**	-1.690**	-2.424***	-1.820**	-1.726**	-1.701**	**-2.680*****	-2.219**	-2.128**	10.781	10.023	**30.614****	**27.873****
5	-1.778**	-1.443*	-1.713**	-2.172**	-1.222*	-1.680**	-2.393***	-2.159**	-2.124**	8.708	8.352	**27.267***	**25.359***
6	-2.224**	-1.584*	-1.332*	-2.277**	-0.974	-1.947**	-2.253**	-2.064**	-1.848**	8.715	8.480	**27.087***	**25.525***
7	-2.267**	-1.613*	-1.332*	-2.413***	-1.034	-1.700**	-2.310**	-2.265**	-2.181**	9.615	9.354	**28.389***	**26.747****
8	-1.941**	-1.882**	-1.146	-2.227**	-0.996	-1.610*	-2.140**	**-2.637****	-2.056**	9.134	8.889	**27.527***	**25.936***
9	-2.092**	-2.019**	-1.500*	-2.212**	-1.053	-1.676**	-2.138**	-2.515**	-2.033**	9.546	9.276	**28.469***	**26.800****
10	-2.086**	-1.793**	-1.291*	-2.267**	-0.716	-1.441*	-1.995**	-2.307**	-1.661**	7.866	7.717	25.479	**24.230***
11	-2.009**	-1.738**	-1.301*	-2.336**	-0.642	-1.311*	-1.921**	-2.122**	-1.956**	7.665	7.535	25.104	**23.936***
12	-1.987**	-1.751**	-1.483*	**-2.730*****	-0.834	-2.102**	-2.024**	-2.332**	-1.643*	9.492	9.308	**28.079***	**26.691****
13	-2.215**	-1.667**	-1.320*	-2.455***	-0.866	-1.846**	-1.710**	-2.050**	-1.549*	7.868	7.673	25.586	**24.159***
14	-2.435***	-1.668**	-1.291*	-2.287**	-0.853	-1.940**	-1.918**	-2.016**	-1.533*	8.218	8.027	**26.126***	**24.715***
15	-2.437***	-1.749**	-1.327*	-2.386***	-1.015	-1.620*	-1.873**	-1.953**	-1.366*	7.933	7.681	25.675	**24.058***
16	-2.301**	-1.591*	-1.102	**-2.639*****	-0.896	-1.387*	-1.543*	-1.967**	-1.366*	7.320	7.114	24.245	22.783

Note: The table reports Augmented Dickey-Fuller (ADF) and and Maddala-Wu (MW) test statistics based on the null hypothesis that the series follows a unit-root process. Due to limitations in the interest rate data, the tests for Belgium, Italy, and Japan are conducted over the period July 1978-December 2004. For all other countries, the tests are conducted over the entire period January 1975-December 2004. All series refer to domestic one-month Eurocurrency deposit interest rates. The maximum number of lags (lags= 16) is selected according to the Schwert (1989) criterion. The critical ADF (332 observations, without trend, without drift) values for rejection of the null hypothesis are given as follows: -2.57 (10% significance level), -2.875 (5%), -3.45 (1%). When allowing for drift, the values of the test statistics stay the same, but now they are compared against the critical values of the Student's t-distribution. The (lower) critical values for two-sided t-test with 316-332 degrees of freedom are given as follows: -1.284 (10% significance level), -1.649 (5%), -2.337 (1%). The last four columns report panel unit-root test statistics calculated according to Maddala and Wu (1999). The results from the panel unit-root tests without drift parameter are reported in the first (entire panel dataset) and the second column (all countries excluding Italy). The last two columns contain results from panel unit-root tests with drift parameter for the entire dataset (third column) and excluding Italy (fourth and last column). Maddala and Wu (1999) panel unit-root test is one sided test that follows χ^2-distribution with $2n$ degrees of freedom (where n is the number of separate time-series). The (upper) critical values for the χ^2-distribution with 18 degrees of freedom (tests for the entire dataset) are given as follows: 25.989 (10% significance level), 28.869 (5%), 34.805 (1%), 42.312 (0.1%). The (upper) critical values for the χ^2-distribution with 16 degrees of freedom (tests that exclude Italy) are given as follows: 23.542 (10% significance level), 26.296 (5%), 32.000 (1%), 39.252 (0.1%).

Bold faces refer to cases where the null hypothesis of nonstationarity (presence of unit root) of the series without allowing for drift is rejected at least at the 10% significance level. When a drift is included in the tests, rejection of the null hypothesis at 10%, 5% and 1% significance level is indicated by *, **, and ***, respectively. In the case of Maddala-Wu panel unit-root tests reported in the last two columns, **bold** faces refer to cases where the null hypothesis of nonstationarity (presence of panel unit-root) is rejected at least at the 10% significance level. Moreover, significance at 10%, 5%, 1%, and 0.1% is indicated by *, **, ***, and ****, respectively.

A.2 Analysis of Total Bias for $\rho \leq 0.9$

Figure A.1 contains a three-dimensional plot from two different angles to describe the range of values for the total bias expression using alternative combinations of ρ and δ. The coefficient of correlation between domestic and foreign interest rates ρ can get any value in the interval $(-1,\ 0.9)$, which is in line with the findings in the actual data.[10] Several interesting observations can be made in this figure. First, the bias can be positive or negative depending on the exact combination of the parameters. Second, the total bias is generally very small in absolute value: its maximum value is about 0.5, while its minimum is about -1.5.[11]

The contour plot given in Figure A.2 describes different value contours of the total bias as a function of the two parameters ρ and δ. Therefore, it shows the precise location of low and high value areas for the total bias within the ρ-δ plane. The highest values, denoted by light contours, are found for combinations of very high ρ and δ close to, but lower than one. Similarly, the lowest values, denoted by dark contours, are found for combinations of very high ρ and δ that are higher than one. Therefore, the general conclusion from Figure A.2 is that the total bias expression obtains extreme values (both positive and negative) when ρ is close to one. Nonetheless, as long as $\rho \leq 0.9$, none of these values is large enough (in absolute value) in order to explain the empirically-found slope coefficient estimates according to equation (3.10).

[10] In fact, only 19 out of 270 "independent" country pairs (or 38 out of 540 pairs in total) have coefficient of correlation that is higher than 0.9.

[11] The values for the relative volatility indicator δ are restricted within the interval (0,2). However, the figure still captures the core characteristics of the total bias function. In fact, the global minimum of the total bias expression almost coincides with the local minimum depicted in the figure.

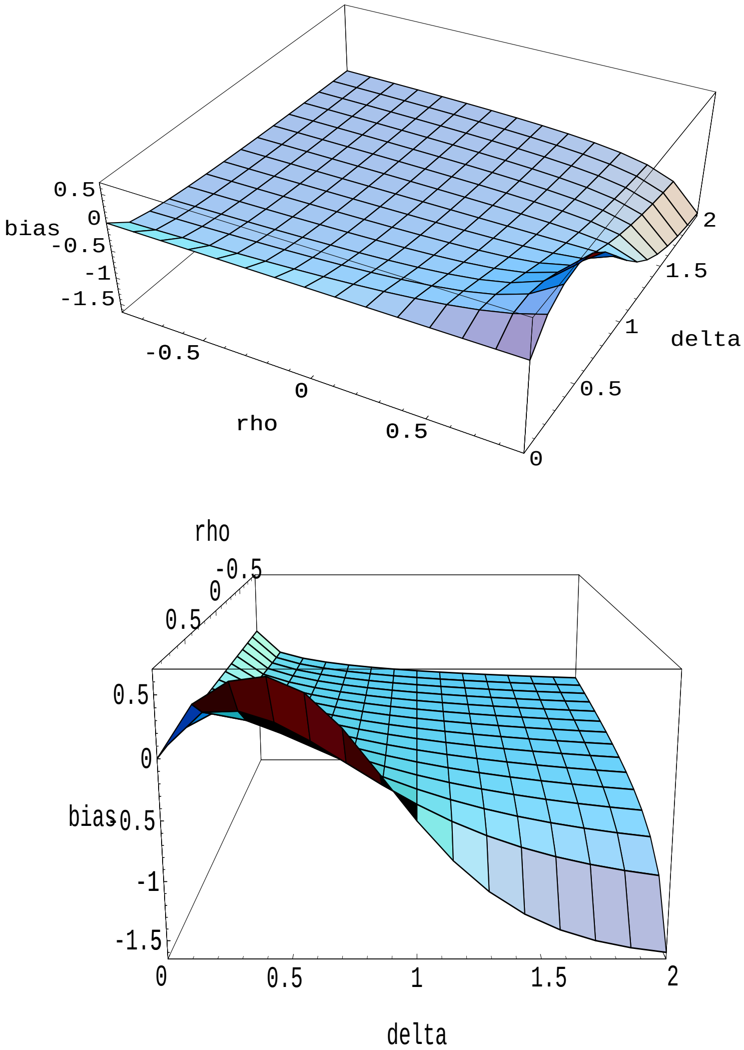


Figure A.1: Total Bias Values for Alternative Combinations of δ and ρ (3D Projections for $\rho \leq 0.9$)

Note: The three-dimensional plots present values for the total bias expression for alternative combinations of δ (relative interest rate volatility) and ρ (coefficient of correlation between the corresponding interest rate series). The relative volatility indicator (delta) can take any value in the interval $(0, 2)$, while the coefficient of correlation can take any value in the interval $(-1, 0.9)$. For convenience and clarity in presentation, we restrict the values for the relative interest rate volatility ($0 \leq \delta \leq 2$). Moreover, the coefficient of correlation between the interest rate series is also restricted ($-1 \leq \rho \leq 0.9$), so that the total bias expression is defined for all combinations of δ and ρ.

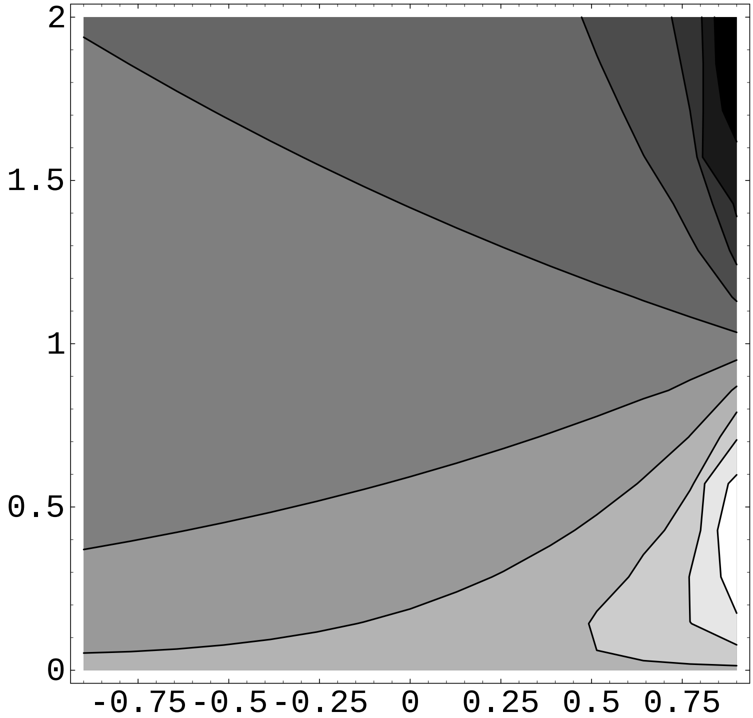


Figure A.2: Total Bias for Alternative Combinations of δ and ρ (contours)

Note: The contour plot presents values for the total bias expression for alternative combinations of δ (relative interest rate volatility) and ρ (coefficient of correlation between the corresponding interest rate series) within two dimensions. Each contour corresponds to a different range of values for the total bias expression. Darker (lighter) contours correspond to lower (higher) values for the total bias expression. The maximal values are depicted by the lightest contour (combination of $\rho = 0.9$ and $\delta < 1$), while the minimal values are depicted by the darkest contour (combination of $\rho = 0.9$ and $\delta > 1$). The relative volatility indicator (delta) can take any value in the interval $(0, 2)$, while the coefficient of correlation can take any value in the interval $(-1, 0.9)$. For convenience and clarity in presentation, we restrict the values for the relative interest rate volatility ($0 \leq \delta \leq 2$). Moreover, the coefficient of correlation between the interest rate series is also restricted ($-1 \leq \rho \leq 0.9$), so that the total bias expression is defined for all combinations of δ and ρ.

Chapter 4

Stochastic Discount Factor Approach to International Risk-Sharing: A Robustness Check of the Bilateral Setting

4.1 Introduction

Depending on the data sources and the theoretical framework used in order to quantify the degree of international risk-sharing, one arrives at very different conclusions. For example, methods that use consumption data and are based on specific underlying utility functions imply that there is not much risk to be shared (consumption growth is not very volatile) and that countries share a very small portion of this risk because cross-country consumption growth correlations are very low (Backus, Kehoe, and Kydland, 1992; Backus and Smith, 1993; Lewis, 1999, 2000; Obstfeld, 1994). Portfolio calculations based on empirical risk-return profiles and certain specification(s) for the utility function find higher potential gains from international risk-sharing (more risk to be shared), but also very low degrees of actual risk diversification (Lewis, 1999, 2000). On the contrary, stochastic discount factor-based measures imply that there is a lot of risk to be shared (high volatility of the discount factors) and that a large portion of this risk is actually shared across countries.

In the latter approach, Brandt, Cochrane, and Santa-Clara (2006) calculate domestic and foreign marginal utility growth rates through stochastic discount factors derived from asset markets data. [1] Subsequently, they compare the volatility of these stochastic discount factors with the volatility of the real exchange rate. Their main finding is that real exchange rates (difference between marginal utility growth rates) are much less volatile than what the stochastic discount factors (proxies for marginal utility growth) of the corresponding countries would imply. Therefore, they conclude that marginal utility growth rates must be very highly correlated across countries, i.e. a large portion of macroeconomic risk is shared internationally.

This chapter presents a robustness check of the (bilateral) stochastic discount factor approach to measuring international risk-sharing given in Brandt, Cochrane, and Santa-Clara (2006). We demonstrate that there are two main limitations of the bilateral SDF approach to international risk-sharing. First, the discount factors in the bilateral framework are not uniquely determined and crucially depend on the partner country included in the calculation. Second, the deviations between the discount factors obtained in this way (the imprecision in the measurement of marginal utility growth) are larger for countries whose stock market excess return shocks are relatively less important (Sharpe ratios are lower).

In order to account for some of these criticisms about the bilateral SDF approach, we extend the bilateral framework into a three-country (trilateral) setting. However, although the trilateral framework demonstrates that the (final) results for the international risk-sharing index are quite robust to the number of countries used in their calculation, it does not resolve the inherent incoherence found in the bilateral SDF model. In fact, it only shifts the problem with the internal incoherence of the SDF approach by one country ahead.

The rest of this chapter is organized as follows: section 4.2 develops the theoretical framework and presents the calculations of the stochastic discount factors and the risk-sharing index. Section 4.3 describes the data, replicates the bilateral results obtained by Brandt, Cochrane, and Santa-Clara (2006),

[1] Following Hansen and Jagannathan (1991), this approach is based on excess returns of the stock market index above the risk-free rate.

and shows some limitations of the bilateral approach. Section 4.4 extends this approach to a three-country setting. We discuss the relevance of our findings in section 4.5. Section 4.6 concludes the chapter.

4.2 Theoretical Framework

4.2.1 Pricing Kernels

In this section we derive the theoretical framework linking the change in the real exchange-rate with the domestic and foreign marginal utility growth rates (stochastic discount factors). Following the approach taken in Backus, Foresi, and Telmer (1996) and Backus, Foresi, and Telmer (2001), we model asset prices with pricing kernels, i.e. stochastic processes that govern the prices of state-contingent securities.[2]

Let v_t represent the domestic currency value at time t of an uncertain, stochastic cash flow of d_{t+1} domestic currency units one period in the future. Then, the basic asset pricing relation relates v_t and d_{t+1} in the following way:

$$v_t = E_t(m_{t+1}d_{t+1}) \tag{4.1}$$

by dividing both sides of equation (4.1) by the initial investment v_t at time t, i.e. the value of the uncertain cash flow at time t, we get an expression in terms of returns:

$$1 = E_t(m_{t+1}R_{t+1}) \tag{4.2}$$

where $R_{t+1} = d_{t+1}/v_t$ is the gross return on this asset/investment between time t and $t+1$, and m_{t+1} is the domestic currency pricing kernel. The kernel m_{t+1} occupies a central place since it gives the "gross rate" at which economic agents discount the uncertain payment d_{t+1} one period in the future, i.e. it

[2]Several conditions should be satisfied in order to derive a relationship between the (real) exchange rate and the stochastic discount factors in the two currencies. First, there should be free trade in assets denominated in each currency as well as free trade in each of the corresponding currencies. Second, no pure (zero initial investment) arbitrage opportunities should exist on any of the markets.

represents the (nominal) intertemporal marginal rate of substitution between time t and $t+1$ for all assets traded in the domestic economy.[3]

Similar relations should hold for assets denominated in foreign currency and traded in the foreign economy. In fact, there are two equivalent ways to show these relations for foreign assets. First, through substitution of all domestic variables from equations (4.1) and (4.2) with their foreign counterparts we get the following equations for foreign assets:

$$v_t^* = E_t(m_{t+1}^* d_{t+1}^*) \tag{4.3}$$

and, in terms of gross returns:

$$1 = E_t(m_{t+1}^* R_{t+1}^*) \tag{4.4}$$

Second, the cash flows (or gross returns) received in foreign currency can be converted into domestic currency units at the expected future spot exchange rate, and then discounted using the domestic pricing kernel or domestic discount factor, just as in the case of domestic assets. According to this approach, we get the following relations:

$$v_t^* = E_t\Big[m_{t+1}(S_{t+1}/S_t)d_{t+1}^*\Big] \tag{4.5}$$

and, in terms of gross returns:

$$1 = E_t\Big[m_{t+1}(S_{t+1}/S_t)R_{t+1}^*\Big] \tag{4.6}$$

where S_t stands for the current spot nominal exchange rate (the price of foreign currency in domestic currency units) at time t, and S_{t+1}/S_t represents its gross rate of change between time t and $t+1$.

Because these two approaches must give equivalent results, we can equate (4.3) with (4.5):

$$E_t(m_{t+1}^* d_{t+1}^*) = E_t\Big[m_{t+1}(S_{t+1}/S_t)d_{t+1}^*\Big] \tag{4.7}$$

[3] m_{t+1} will be a unique solution of equations (4.1) and (4.2) only if the domestic economy has a complete set of state-contingent securities that can be freely traded. Otherwise, there are multiple solutions for m_{t+1}.

or (4.4) with (4.6), respectively:

$$E_t(m^*_{t+1}R^*_{t+1}) = E_t\Big[m_{t+1}(S_{t+1}/S_t)R^*_{t+1}\Big] \tag{4.8}$$

If no pure arbitrage opportunities exist and markets in both countries are complete, then the following should hold:[4]

$$m^*_{t+1} = m_{t+1}(S_{t+1}/S_t) \tag{4.9}$$

which, in turn, gives the relation between the change of the exchange rate and the nominal discount factors in the two countries. Hence, the (nominal) exchange rate should move (depreciate/appreciate) exactly by the difference between the discount factors in the respective countries. More specifically, equation (4.9) implies that domestic currency depreciates when the domestic nominal discount factor is lower than the foreign nominal discount factor in the corresponding period.

Although the discussion in this section focused on *nominal* variables, a similar condition can be stated in terms of *real* variables. Thus, taking the logarithm of both sides of equation (4.9) and changing all nominal variables (exchange rates, gross returns, discount factors) into their real counterparts, we arrive at a condition that equates the real exchange rate to the difference between changes in foreign and domestic intertemporal marginal rates of substitution between time t and $t+1$:

$$\ln\frac{e_{t+1}}{e_t} = \ln\frac{\lambda^*_{t+1}}{\lambda_{t+1}} = \ln\lambda^*_{t+1} - \ln\lambda_{t+1} \tag{4.10}$$

where e_t is the real exchange rate - the relative price of foreign in terms of domestic goods,[5] λ_{t+1} is the gross rate of change in domestic marginal utility between time t and $t+1$, λ^*_{t+1} is the gross rate of change in foreign marginal utility between time t and $t+1$ (both measured in units of real,

[4]This relation holds in the case of complete markets in both countries (for currencies and risky assets). In incomplete markets, m^*_{t+1} and m_{t+1} will not be uniquely determined - combinations of the discount factors with some random disturbances ϵ^*_{t+1} and ϵ_{t+1} that are orthogonal to the underlying shocks will also price all assets.

[5]The real exchange rate is defined as the price of foreign goods over the price of domestic goods. Therefore, an increase in the real exchange rate implies a real appreciation (depreciation) of foreign (domestic) goods.

consumption goods).[6] Rearranged in real terms, this condition states that in equilibrium the change in the relative price of foreign in terms of domestic goods (given by gross rate of change in the real exchange rate) should equal the ratio between foreign and domestic marginal utility changes (stochastic discount factors or pricing kernels). Derived through this simple asset pricing framework, equation (4.10) is of central importance for the stochastic discount factor approach to measuring international risk-sharing, elaborated in this study.[7]

4.2.2 Risk-Sharing Index

The perfect international risk-sharing hypothesis implies complete equalization of marginal utility growth rates across countries. In our framework, given by equation (4.10), it means equality between λ_{t+1} and λ^*_{t+1} at any point in time. Thus, if this asset pricing condition holds *and* all country-specific risks are shared internationally, then the left-hand side of this equation should always be zero. Put differently, the departures from this perfect situation can be measured by the deviations on the left-hand side, i.e. the fluctuations of the real exchange rate.

Brandt et al. (2006) use this intuition to propose a measure of international risk-sharing based on asset markets. First, they take variances of both sides of equation (4.10):

$$\sigma^2\Big(\ln\frac{e_{t+1}}{e_t}\Big) = \sigma^2\Big(\ln\lambda^*_{t+1} - \ln\lambda_{t+1}\Big) = \\ = \sigma^2\Big(\ln\lambda^*_{t+1}\Big) + \sigma^2\Big(\ln\lambda_{t+1}\Big) - 2\rho\sigma\Big(\ln\lambda^*_{t+1}\Big)\sigma\Big(\ln\lambda_{t+1}\Big) \quad (4.11)$$

where σ^2 symbolizes a variance, σ a standard deviation, and ρ is the coefficient of correlation between the two discount factors λ_{t+1} and λ^*_{t+1}. There-

[6] The stochastic discount factors λ_{t+1} and λ^*_{t+1} represent gross real returns in the corresponding markets. They can be defined through in traditional consumption-based models as $\lambda_{t+1} = \beta(u'(c_{t+1}/u'(c_t))$, where β is the reciprocal of the gross rate of time preference and $(u'(c_{t+1}/u'(c_t))$ is the gross rate of change in marginal utility growth between time t and $t+1$. Therefore, the values for the discount factors will be always positive in this framework, typically in the vicinity of 1.

[7] For more extensive discussion on the application of this equation see Backus et al. (2001) and Brandt and Santa-Clara (2002) for example.

fore, if the following two conditions hold: i) assets and currencies are priced according to equation (4.10) at any point in time; and ii) all risks are shared internationally, then: $\rho = 1$, $\lambda_{t+1} = \lambda^*_{t+1}$ and $\sigma^2\Big(\ln \frac{e_{t+1}}{e_t}\Big) = 0$. In general, the correlation between marginal utility growth rates will be given by:

$$\rho = \frac{\left[\sigma^2\Big(\ln \lambda^*_{t+1}\Big) + \sigma^2\Big(\ln \lambda_{t+1}\Big) - \sigma^2\Big(\ln \frac{e_{t+1}}{e_t}\Big)\right]}{2\sigma\Big(\ln \lambda^*_{t+1}\Big)\sigma\Big(\ln \lambda_{t+1}\Big)} \tag{4.12}$$

indicating that risk-sharing across countries decreases in the variability of the real exchange rate. Based on this idea, Brandt et al. (2006) construct the following risk-sharing index

$$RSI = 1 - \frac{\sigma^2\Big(\ln \frac{e_{t+1}}{e_t}\Big)}{\sigma^2\Big(\ln \lambda^*_{t+1}\Big) + \sigma^2\Big(\ln \lambda_{t+1}\Big)} \tag{4.13}$$

where the numerator of the second term captures the variability in the real exchange rate (which, according to the argumentation above, measures the deviations from perfect risk-sharing), and the denominator is the sum of the variabilities in marginal utility growth in the two countries (the total risk that exists and can be shared across countries). Hence, this term gives a ratio between risk still not shared and total risk that can be shared between the two countries. Brandt et al. (2006) indicate that this index gives the portion of total (diversifiable) risk that is already shared by the two countries.[8]

4.2.3 Basic Calculations

In order to calculate the risk-sharing index given in the previous section, first we have to recover the log discount factors (or marginal utility growth rates) from asset markets data in the corresponding countries.[9] For this purpose, we closely follow the exposition given in Brandt et al. (2006). We start by assuming that the following assets are traded in a two-country setting:

[8] In this way, the framework presented by Brandt et al. (2006) can be viewed as an extension of the Hansen-Jagannathan (1991) volatility bounds to the international setting.

[9] For ease of exposition and manipulation in the further calculations (translating between levels and logarithms), the demonstration here uses continuous time formulation. Empirically, all variables are calculated using the corresponding discrete time approximations, see the section on data issues.

$$\frac{dB^d}{B^d} = r^d dt \tag{4.14}$$

$$\frac{dS^d}{S^d} = \theta^d dt + dz^d \tag{4.15}$$

$$\frac{de}{e} = \theta^e dt + dz^e \tag{4.16}$$

$$\frac{dB^f}{B^f} = r^f dt \tag{4.17}$$

$$\frac{dS^f}{S^f} = \theta^f dt + dz^f \tag{4.18}$$

where B^d is the domestic risk-free bond (with expected return r^d), S^d is the domestic risky asset (expected return θ^d), e is the real exchange rate, i.e. the relative price of foreign in terms of domestic goods (expected return θ^e), B^f is the foreign risk-free bond, and S^f is the foreign risky asset (expected return θ^f). There are three sources of uncertainty in this setting, related to the domestic asset, the real exchange rate, and the foreign asset. These shocks can be collected into a vector of shocks dz:

$$dz = \begin{bmatrix} dz^d \\ dz^e \\ dz^f \end{bmatrix}$$

with a corresponding variance-covariance matrix given by:[10]

$$\Sigma = \frac{1}{dt}E(dzdz') = \begin{bmatrix} \Sigma^{dd'} & \Sigma^{de} & \Sigma^{df'} \\ \Sigma^{ed'} & \Sigma^{ee} & \Sigma^{ef'} \\ \Sigma^{fd'} & \Sigma^{fe} & \Sigma^{ff'} \end{bmatrix}$$

Furthermore, the calculation of the discount factor(s) from asset markets depends primarily on the variability of the excess returns on risky assets, driven by the shocks in vector dz.[11] We derive all excess return equations in

[10] This variance-covariance matrix is the same for domestic and foreign investors because they face the same vector of shocks in this symmetric, bilateral setting.

[11] Since we work with (expected) excess returns in this analysis, we do not make a real/nominal returns distinction.

the appendix, and here present only their expected values. Thus, the domestic investor faces the following set of expected excess returns:

$$\mu^d = \begin{bmatrix} \theta^d - r^d \\ \theta^e + r^f - r^d \\ \theta^f - r^f + \Sigma^{ef} \end{bmatrix}$$

The first term in this vector gives the excess return that a domestic resident expects to get by investing on the domestic stock market. It equals the difference between the average real return on the domestic stock market index (θ^d) and the average real risk-free rate in the domestic economy (r^d) during the entire investment period. The expected excess return on the foreign exchange market is given by the second term in vector μ^d. It represents the average deviation from (uncovered) interest parity, calculated as borrowing in the domestic currency, converting the borrowed amount into the foreign currency, lending at the ongoing one-month foreign interest rate, and converting the proceeds back into domestic currency after one month. The last term in vector μ^d gives the expected excess return that a domestic investor expects to get by investing in the foreign stock market. Therefore, it represents a difference between the average return on the foreign stock market and the domestic one-month risk-free interest rate. The last part of this term Σ^{ef} results from the continuous-time formulation and gives the (average) co-movement between the returns on the foreign stock market and the exchange rate. Therefore, by correcting for the movements of the nominal exchange rate, this term facilitates the translation of excess returns obtained on the foreign market.[12]

A similar vector of expected excess returns applies to the foreign investor:

$$\mu^f = \begin{bmatrix} \theta^d - r^d - \Sigma^{ed} \\ -(\theta^e + r^f - r^d - \Sigma^{ee}) \\ \theta^f - r^f \end{bmatrix}$$

The interpretation of the terms is analogous to that given for the domestic investor. The expected excess return on the foreign exchange market is exactly

[12]For example, Σ^{ef} is added to the excess return on the foreign market for the domestic investor, suggesting that foreign expected excess returns are amplified when associated with appreciation of the foreign currency.

the opposite of the one for the domestic investor (corrected for the continuous-time term Σ^{ee}).

Then, the following discount factors price all assets according to the basic pricing conditions:[13]

$$\frac{d\Lambda^i}{\Lambda^i} = -r^i dt - \mu^{i\prime}\Sigma^{-1}dz, i = d, f \tag{4.19}$$

where $\frac{d\Lambda^i}{\Lambda^i}$ is the growth rate of the discount factor, r^i is the risk-free return, and μ^i is the vector of excess returns for risky assets in country i. In order to calculate the change in the log discount factor $\ln\lambda^i$ required in equation (4.10), we use Ito's lemma and get the following expression:

$$d\ln\Lambda^i = \frac{d\Lambda^i}{\Lambda^i} - \frac{1}{2}\frac{d\Lambda^{i2}}{\Lambda^{i2}} = -\Big(r^i + \frac{1}{2}\mu^{i\prime}\Sigma^{-1}\mu^i\Big)dt - \mu^{i\prime}\Sigma^{-1}dz \tag{4.20}$$

and for its standard deviation:

$$\frac{1}{dt}\sigma^2(d\ln\Lambda^i) = \mu^{i\prime}\Sigma^{-1}\mu^i, i = d, f \tag{4.21}$$

The change in the log discount factor $d\ln\Lambda$ corresponds to $\ln\lambda_{t+1}$ in the basic asset pricing condition (4.10). Therefore, the risk-sharing index given by (4.13) can be calculated directly from the second moments according to the following expression:

$$RSI = 1 - \frac{\sigma^2(d\ln\Lambda^d - d\ln\Lambda^f)}{\sigma^2(d\ln\Lambda^d) + \sigma(d\ln\Lambda^f)} = 1 - \frac{\Sigma^{ee}}{\mu^{d\prime}\Sigma^{-1}\mu^d + \mu^{f\prime}\Sigma^{-1}\mu^f} \tag{4.22}$$

In order to show the symmetric structure of our framework, we relate the shocks facing the domestic with those facing the foreign investor. The expected excess returns vectors μ^d and μ^f differ only by the exchange rate changes:[14]

$$\mu^d - \mu^f = \begin{bmatrix} \theta^d - r^d \\ \theta^e + r^f - r^d \\ \theta^f - r^f + \Sigma^{ef} \end{bmatrix} - \begin{bmatrix} \theta^d - r^d - \Sigma^{ed} \\ \theta^e + r^f - r^d - \Sigma^{ee} \\ \theta^f - r^f \end{bmatrix} = \begin{bmatrix} \Sigma^{ed} \\ \Sigma^{ee} \\ \Sigma^{ef} \end{bmatrix} \tag{4.23}$$

[13] For more details on finding the discount factor in this setting see Brandt et al. (2006, p.675-677) or Chapter 4 in Cochrane (2004).

[14] In order to derive this relation, we disregard the change in sign before the foreign exchange excess returns when moving from domestic to foreign investor perspective.

From these formulae, it is clear that the expected excess return vectors differ exactly by the middle column of the common variance covariance matrix Σ^e:

$$\mu^d - \mu^f = \begin{bmatrix} \Sigma^{ed} \\ \Sigma^{ee} \\ \Sigma^{ef} \end{bmatrix} = \Sigma^e \tag{4.24}$$

In turn, we can derive a relationship between the domestic and foreign discount factor loadings (given by the last term of equation (4.20)):

$$\mu^d \Sigma^{-1} = (\mu^f + \Sigma^e)\Sigma^{-1} = \mu^f \Sigma^{-1} + \Sigma^e \Sigma^{-1} = \mu^f \Sigma^{-1} + \begin{bmatrix} 0 \\ 1 \\ 0 \end{bmatrix} \tag{4.25}$$

Equation (4.25) shows that domestic and foreign discount factors load equally on domestic and foreign stock market shocks, while their loadings on the foreign exchange shocks differ by exactly 1. Therefore, this implies that the only difference between the two discount factors comes from fluctuations in the real exchange rate.

4.3 Data and Replication of Results

4.3.1 Data Description

In this section we replicate the results for the bilateral setting presented in Brandt et al. (2006). For that purpose, we construct a dataset that is as close as possible to the one used in the original study. In particular, we employ three types of time-series: for the risk-free rate we use interest rates on one-month Eurocurrency deposits, while for the return on the risky asset we use total returns on the stock market index for the corresponding country. We calculate inflation rates from the changes in the consumer price indices (CPI). The nominal exchange rates are expressed in terms of domestic currency per unit of foreign currency.

Our analysis includes three economies: USA, UK, and Japan. We use monthly data from January 1975 till June 1998 for the USA and the UK.

For Japan interest rates on Eurocurrency deposits are not available before August 1978. Therefore, all data series for Japan start in August 1978 and go through June 1998. The series on Eurocurrency deposit interest rates, nominal exchange rates and total stock market index returns are measured at the beginning of the month, while the CPI series refer to mid-month values.[15] All data come from Datastream.[16]

For stock market returns, we use the same indices employed in the original study:[17] S&P 500 for the USA, FTSE ALL for the UK, and NIKKEI 225 for Japan.

4.3.2 Summary Statistics

We use discrete time approximations of the continuous time formulae derived in section 4.2.3. The following sample counterparts are used in the calculation:

$$\begin{aligned}
&\theta^d - r^d = \tfrac{1}{\Delta} E_T R^d_{t+\Delta} \\
&\theta^f - r^f = \tfrac{1}{\Delta} E_T R^f_{t+\Delta} \\
&\theta^e + r^f - r^d = \tfrac{1}{\Delta} E_T \left(\tfrac{e_{t+\Delta} - e_t}{e_t} + r^f_{t+\Delta} - r^d_{t+\Delta} \right) \\
&dz^d = \tfrac{1}{\Delta} (R^d_{t+\Delta} - E_T R^d_{t+\Delta}) \\
&dz^f = \tfrac{1}{\Delta} (R^f - E_T R^f_{t+\Delta}) \\
&dz^e = \tfrac{1}{\Delta} \left(\tfrac{e_{t+\Delta} - e_t}{e_t} \right) - \tfrac{1}{\Delta} E_T \left(\tfrac{e_{t+\Delta} - e_t}{e_t} \right) \\
&\Sigma = E_T(dz dz')
\end{aligned}$$

In these sample moments T is the sample size (281 monthly observations), E_T denotes the sample mean for the entire time period, $\Delta = \frac{1}{12}$years, $R^d_{t+\Delta}$ and $R^f_{t+\Delta}$ correspond to the domestic and foreign excess stock returns, and $r^d_{t+\Delta}$ and $r^f_{t+\Delta}$ refer to the domestic and foreign risk-free (Eurocurrency deposits) interest rates, respectively.[18]

[15] The results are very robust with respect to the use of lag or lead values for the inflation rate

[16] CPI data is retrieved from Datastream and comes from the IMF International Financial Statistics (IFS) database.

[17] For the UK we do the same calculations using FTSE 100 index. The results change only slightly.

[18] The formulae for expected excess returns and shocks on the stock markets and the foreign exchange market are annualized through division by $\Delta = \frac{1}{12}$years.

In accordance with the approach taken before, we use real variables: real (excess) stock returns, real risk-free interest rates and real exchange rates. Hence, we correct all data series by the inflation rate (measured by changes in the mid-month CPI).[19] Moreover, we calculate stock market returns in two ways: i) assuming continuous-time specification and ii) with discrete time specification. Since the results are very similar, in the rest of the analysis we only present stock market returns calculated using the discrete time framework.

The summary statistics are presented in Table 4.1. Its upper panel shows means and standard deviations for excess stock market returns (Stock) and for excess foreign exchange returns (X-rate). The former are derived as returns on the stock market index above the one-month Eurocurrency interest rate, while the latter are derived as deviations from the uncovered interest parity (UIP), calculated as excess returns from borrowing in the domestic currency (dollar), investing in one-month Eurocurrency deposits in the foreign country (pounds sterling or yen), and translating these yields back to the domestic currency at the end of the period. All entries in the table are annualized and reported in percentages.

The statistics in Table 4.1 are very similar to and convey the same message as the ones presented by Brandt et al. (2006).[20] In fact, the mean excess returns given in the first row illustrate the high equity premium found in stock markets data. They range from 4.29 percent in Japan, 9.97 percent in the USA, to 10.21 percent in the UK. All of them are statistically different from zero. Moreover, their associated standard errors, reported in the row beneath, are typically very high. Thus, they result in values for the Sharpe ratio between 0.22 for Japan, 0.62 for the UK, to 0.72 for the USA. Therefore, these results suggest that investors in the USA got the highest excess returns per unit of risk taken, while investors in Japan got the lowest. On the other hand, mean excess returns for foreign exchange are much smaller and not statistically

[19] Our main results are based on excess market returns. Therefore, they are not sensitive to whether nominal or real variables are used in the calculations.

[20] The first moments are similar and normally keep the same ranking between different countries, but are not identical. On the other hand, the second moments are almost identical as the ones presented by Brandt et al. (2006). This is to be expected as the second moments are usually much less sensitive to the exact procedure used in the calculation.

Table 4.1: Summary Statistics (Annualized)

	USA	UK		Japan		
	Stock	Stock	X-Rate ($\$/£$)	Stock	X-Rate ($\$/Y$)	X-Rate ($£/Y$)
			Returns (%)			
Mean	9.97	10.21	0.98	4.29	2.06	1.08
Std Dev	13.80	16.49	11.77	19.52	12.67	12.16
Sharpe ratio	0.72	0.62	0.08	0.22	0.16	0.09
	USA	UK		Japan		
	Stock	Stock	X-Rate ($\$/£$)	Stock	X-Rate ($\$/Y$)	X-Rate ($£/Y$)
			Return Correlations			
USA Stock	1					
UK Stock	0.583	1				
X-Rate ($\$/£$)	0.010	-0.050	1			
Japan Stock	0.324	0.342	0.077	1		
X-Rate ($\$/Y$)	-0.023	-0.063	0.507	0.101	1	
X-Rate ($£/Y$)	-0.037	0.065	-0.439	0.030	0.551	1

Note: The table contains summary statistics and correlations for real excess returns on stock and foreign exchange markets. All figures are calculated over the time period January 1975-June 1998 (for USA and UK) or over the period August 1978-June 1998 (for Japan). The upper panel figures for the means, standard deviations and Sharpe ratios of all shocks. The lower panel contains figures for the coefficient of correlation between the corresponding returns. Stock market excess returns are calculated as returns on the stock market indices over the one-month Eurocurrency deposit rate for the corresponding country/currency. Excess returns on the foreign exchange market are calculated as (real) deviations from uncovered interest rate parity ($\theta^e + r^f - r^d$): borrowing at the US interest rate, converting to the foreign currency, investing on the foreign interest rate, and converting the proceeds back to US dollars. All data-series are retrieved from Datastream. The summary statistics presented in the upper panel are annualized and expressed in percentage terms (rounded to two decimal places).

different from zero.[21] Furthermore, the annualized standard deviations for foreign exchange excess returns are about half the values for excess stock market returns (11.56 percent for the first, 12.67 percent for the second, and 12.16 for the third exchange rate).

Finally, the lower panel of this table presents a returns correlation matrix. Three conclusions are evident from this table. First, foreign exchange excess returns are very weakly correlated with excess returns on stock markets. Second, foreign exchange excess returns on one currency pair are highly correlated with excess return on the other currency pair (correlations of 0.507. 0.551 and -0.439). Third, excess returns for different stock markets are highly correlated among themselves (correlations ranging from 0.32 between USA and Japan to 0.58 between USA and UK).

[21] In fact, all mean excess returns on the foreign exchange market are within the range 1-2 percent.

4.3.3 Replication of the Results for the Bilateral Setting

Results for the Risk-Sharing Index

Using the dataset described in the previous section, here we present a replication of the results obtained by Brandt et al. (2006) for the bilateral setting. The most important result is presented in the first row of Table 4.2. The risk-sharing index obtains values higher than 0.98, which indicates that an extremely large portion of total macroeconomic risks faced by investors in different countries is shared internationally. This is the central result and the most important message from Brandt et al. (2006). In order to understand these high values for the risk-sharing index, we present its two components in the lower part of Table 4.2. The volatility of the real exchange rate (numerator in the second term of the risk-sharing index) is several times lower than the volatility of the stochastic discount factors, i.e. the volatility of the intertemporal marginal utility growth rates (denominator in the second term of the risk-sharing index). In fact, the discount factors calculated from asset markets are very volatile, implying that marginal utility varies by about $65-75$ percent per year.[22] In turn, this implies low values for the second term in equation (4.13) and high value for the overall risk-sharing index.

Discount Factor Loadings

The volatility of the stochastic discount factor (marginal utility growth rate) comes from three sources: domestic and foreign stock market excess return shocks and the foreign exchange excess return shock. The loadings on each of these shocks enter the equations for the discount factors with a negative sign, meaning that a positive shock leads to a decrease in the discount factor (equation (4.19)). For example, a positive (negative) shock on the US stock market (dz^d) leads to a decrease (increase) in domestic and foreign marginal utility growth rates (discount factor levels).[23] Table 4.3 presents figures for

[22]The volatility of the stochastic discount factor crucially depends on the (average) excess returns earned by the asset markets (equation (4.21)). Therefore, high values for the discount factor volatility reflect the (abnormally) high equity premium earned by investors (Mehra and Prescott, 1985; Kocherlakota, 1996).

[23]A favorable stock market shock leads to lower marginal utility growth rate as shown by the negative sign in front of the disturbance term in equation (4.20). Moreover, this shock

Table 4.2: Risk Sharing Index

	USA vs. UK	USA vs. Japan	UK vs. Japan
Risk Sharing Index	0.9878	0.9857	0.9821
Real X-Rate Volatility	11.75	12.47	12.05
Volatility of Marginal Utility Growth:			
Domestic	75.49	74.83	62.21
Foreign	75.11	73.09	65.06

Note: The table presents results for the bilateral risk-sharing index. The first row gives figures for the overall risk-sharing index calculated according to the following formula: $RSI = 1 - \frac{\Sigma^{ee}}{\mu^{d\prime}\Sigma^{-1}\mu^{d}+\mu^{f\prime}\Sigma^{-1}\mu^{f}}$. The second row refers to the volatility of the real exchange rate found in the numerator of the risk-sharing index, while the last two rows refer to the volatility of the stochastic discount factors found in the denominator of the risk-sharing index. Domestic refers to the first country, while foreign refers to the second country mentioned in the country-pair. The volatilities of the real exchange rate and the marginal utility growth are measured as annualized standard deviations and are expressed in percentage terms (rounded to two decimal places).

the discount factors loadings ($\mu^d\Sigma^{-1}$ and $\mu^f\Sigma^{-1}$) on each of these underlying shocks.

In line with equation (4.25), domestic and foreign discount factors are restricted to load equally on each of the stock market shocks, and the domestic discount factor loads on the exchange rate shocks by one more than the foreign discount factor. The last point implies that the difference between the two discount factors at each point in time equals the fluctuations in the real exchange rate.[24] Furthermore, these foreign exchange loadings are of similar magnitude in all three country-pairs (in absolute value terms) and are always lower than the dominant stock market loadings.

There are large differences between stock markets discount factor loadings for each of the three bilateral country-pairs. For example, the loadings on the domestic (USA) stock market (3.76 and 5.36, respectively) are much higher than the loadings on the other two stock markets (1.95 for UK and -0.13 for Japan) in the first two country-pairs. This suggests that the USA stock market represents the dominant source of variability for both discount factors (domestic and foreign) for these pairs (USA vs. UK and USA vs. Japan). In fact, this finding reflects the superior return compensation per unit of risk

is "scaled" by the loading coefficient $\mu'\Sigma^{-1}$.

[24]This reflects the symmetric nature of the foreign exchange excess return shocks given by equation (4.25).

Table 4.3: Discount Factor Loadings (Bilateral)

	USA vs. UK		USA vs. Japan		UK vs. Japan	
	Domestic	Foreign	Domestic	Foreign	Domestic	Foreign
dz^d	3.76	3.76	5.36	5.36	3.69	3.69
dz^e	-1.02	-2.02	1.51	0.51	-0.41	-1.41
dz^f	1.95	1.95	-0.13	-0.13	0.12	0.12

Note: The table presents figures for the discount factor loadings in the bilateral setting. The loadings for the domestic discount factor are given by $\mu^{d\prime}\Sigma^{-1}$ and the corresponding loadings for the foreign discount factor are given by $\mu^{f\prime}\Sigma^{-1}$. For each of the three bilateral country-pairs domestic refers to the first country and foreign refers to the second country mentioned in the country-pair. The row marked dz^d contains figures for discount factor loadings on the domestic stock market shocks, row dz^e refers to discount factor loadings on the foreign exchange market shocks, and row dz^f refers to discount factor loadings on the foreign stock market shock for the corresponding country-pair.

undertaken that investors get in the USA compared to the other two stock markets given by the Sharpe ratios in Table 4.1. Since investors' utility directly depends on the Sharpe ratio, i.e. the compensation they get per unit of risk, excess return shocks on markets/assets with the highest Sharpe ratio matter more for the stochastic discount factor (marginal utility growth). Therefore, excess return shocks on the USA stock market matter most, while shocks on the Japanese stock market matter the least for investors' utility changes.

Furthermore, the discount factors load negatively (and load much less in absolute value) on the Japanese excess return shocks in the second country-pair (USA vs. Japan). This finding (partially) reflects the low price of risk on the Japanese relative to the American stock market (Sharpe ratio of 0.22 for Japan compared to 0.72 for the USA). In fact, since the Japanese stock market is clearly dominated by the American stock market, holding any non-negative investment position on the Japanese market implies that investors forego better investment opportunities on the American market. Hence, this sub-optimal behavior explains the anomalous loadings on the Japanese stock market reported in the middle columns of Table 4.3.

Visual Evidence

In order to give a visual representation of the main result in our study, we present several plots for the discount factors. First, in Figure 4.1 we show time paths for the log discount factors in the three country pairs. We calculate the log level of the discount factor in line with equation (4.20). It contains two components: a trend component given by the expected value of equation (4.20) (the term in brackets) and a disturbance component given by the loadings on the underlying excess return shocks. The development of the log level discount factors can be best understood through the contribution of each of its components.

There are several interesting issues in this figure. First, the log level discount factors typically slope downward as a result of the trend component. In fact, as long as the sum of the average real risk-free rate and the discount factor volatility (the expected value of equation (4.20) given by the term in brackets) is positive (as normally observed), the log level discount factors will follow a downward trend. The easiest way to understand why this is usually the case is by looking at an economy with one only risk-free bond. If this economy experiences real growth over an extended period of time, then its average real risk-free interest rate will be positive (and the trend component will be negative). That is, a downward trend in the log level discount factor corresponds with a decreasing trend in marginal utility growth rates or continual improvement in overall economic conditions. Second, it is clear from the figure that both discount factors follow a similar pattern and move very closely together. In fact, the only difference between them comes from the real exchange rate fluctuations (see equations (4.10) and (4.25)). Based on this observation, we can conclude that marginal utility growth rates across countries follow very similar time paths, just as implied by the perfect risk-sharing condition.

Moreover, in Figure 4.2 we present scatterplots for the discount factor growth rates. We calculate these monthly growth rates according to equation (4.19). This figure just strengthens our conclusion from Figure 4.1 : there is a very high positive correlation between the discount factor growth rates for each country pair. Most observations/points are literally lying on the 45 degree line, thereby indicating that the stochastic discount factor approach implies nearly perfect levels of (bilateral) international risk-sharing.

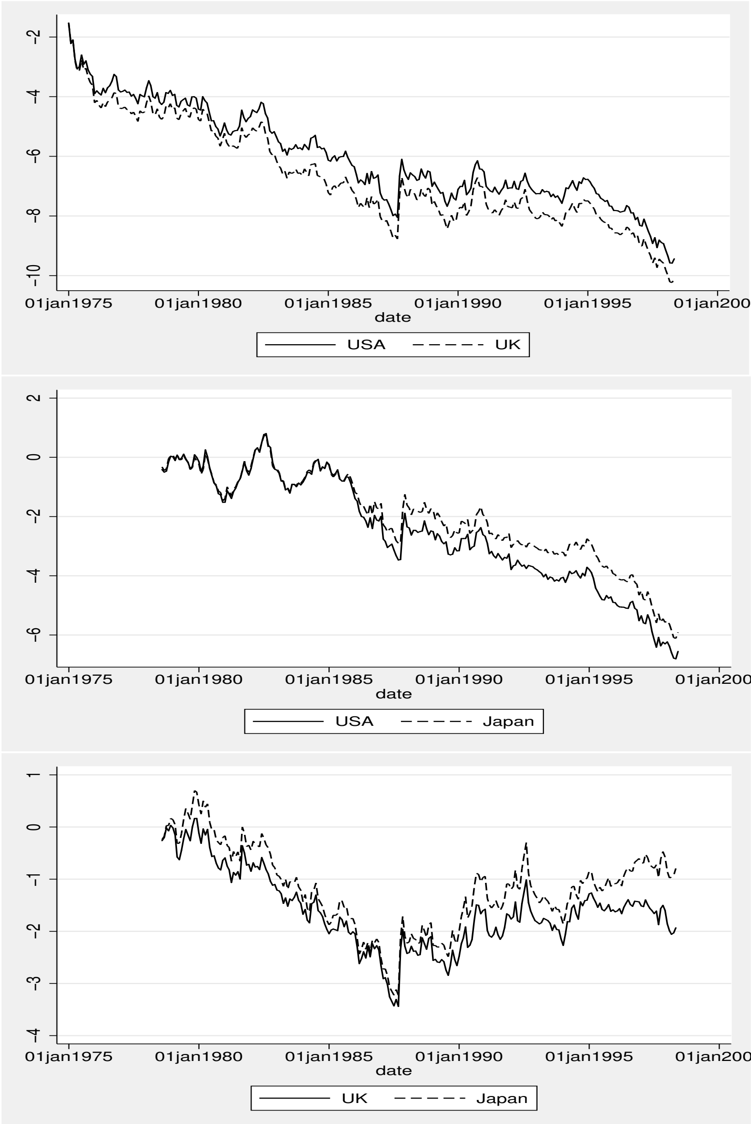


Figure 4.1: Log Levels of Discount Factors (Bilateral)

Note: The figure presents time lines of the log levels of the discount factors calculated in the bilateral setting. Each plot refers to separate country-pair. The log levels of the discount factors are calculated through accumulation of the changes in the log discount factors given in equation (4.20).

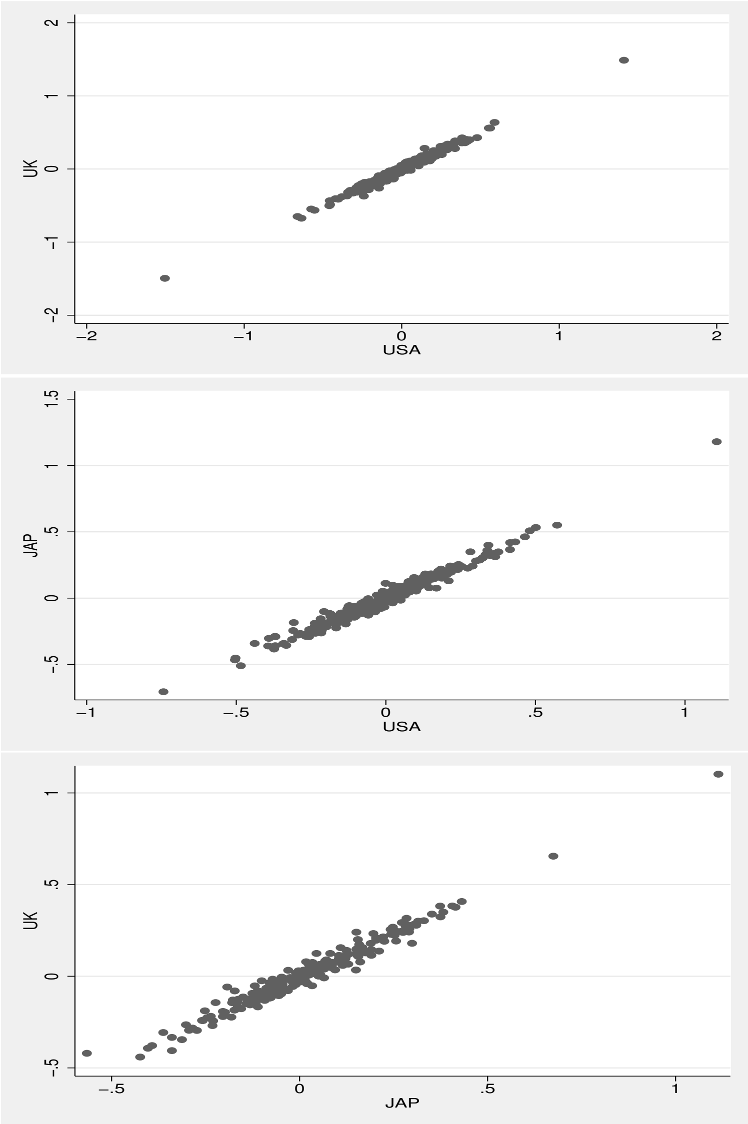


Figure 4.2: Growth of Discount Factors (Bilateral)

Note: The figure presents scatterplots for growth rates of the discount factors calculated in the bilateral setting. Each plot refers to separate country-pair. The growth of discount factors is calculated according to equation (4.19).

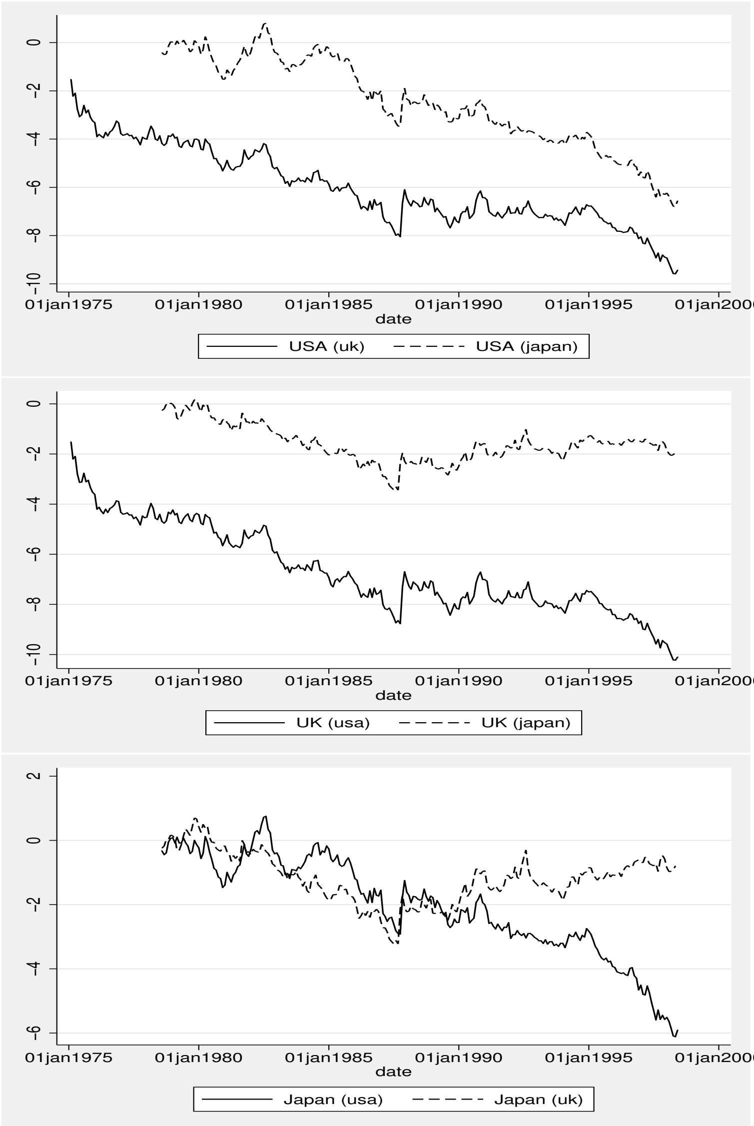


Figure 4.3: Comparison of Log Levels of Discount Factors (Bilateral)

Note: The figure presents time lines of the log levels of the discount factors calculated in the bilateral setting. Each plot refers to log levels for one country when alternative countries are used as partners. The log levels of the discount factors are calculated through accumulation of the changes in the log discount factors given in equation (4.20).

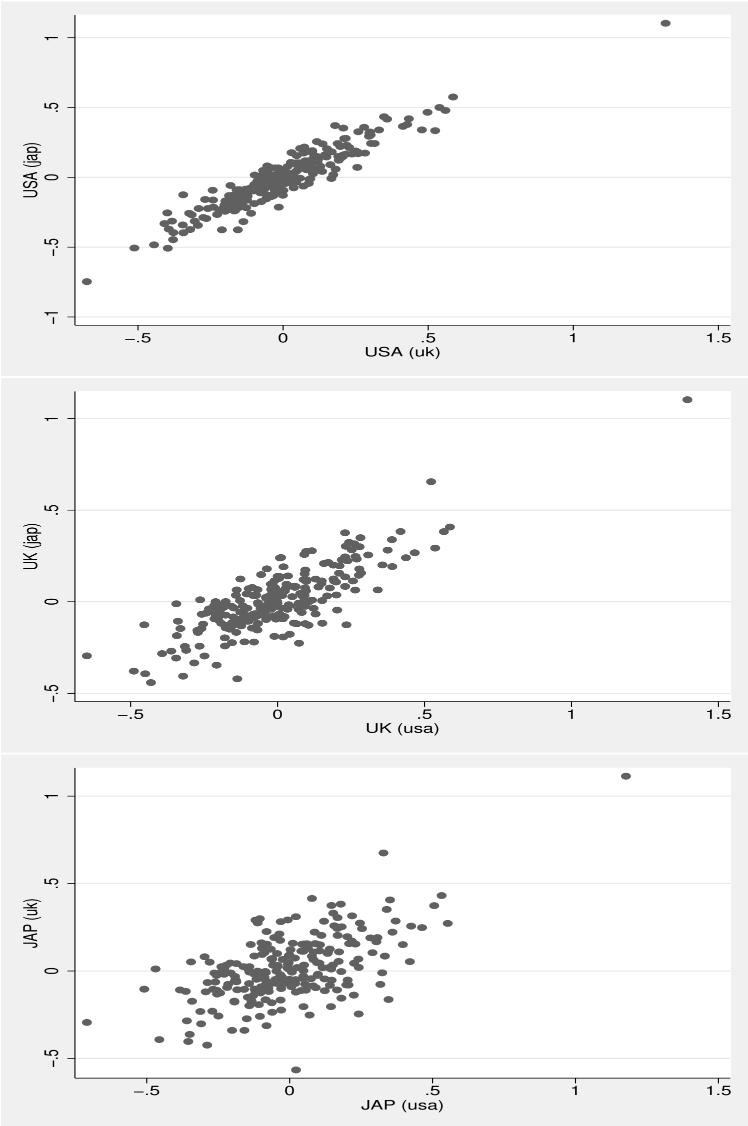


Figure 4.4: Comparison of Discount Factor Growth Rates (Bilateral)

Note: The figure presents scatterplots for growth rates of the discount factors calculated in the bilateral setting. Each plot refers to discount factor growth for one country when alternative countries are used as partners. The growth of discount factors is calculated according to equation (4.19).

4.3.4 Discussion about the Results from the Bilateral Setting

Section 4.3.3 demonstrated that measures based on the stochastic discount factor approach imply very high levels of international risk-sharing among three different country-pairs: USA-UK, USA-Japan, and UK-Japan. In fact, we showed that discount factors for each country in the bilateral pair display very similar levels of volatility (Table 4.2), follow similar time paths (Figure 4.1), and have almost identical growth rates (Figure 4.2). However, all these calculations were conducted within a bilateral setting, i.e. treating only two countries at the time. Therefore, one possible criticism of this approach is that a country's discount factor obviously depends on the choice of the second country. In particular, the USA log discount factor displays a very similar behavior with the UK log discount factor (in the first panel of Figure 4.1). Similarly, in the second panel of Figure 4.1, the USA and Japan discount factors are much alike too. However, the USA log discount factor from the first panel is quite different from the USA log discount factor given in the second panel. Correspondingly, the difference between the two UK discount factors in the first and the third panel and between the two Japan discount factors in the second and the third panel is even larger. In other words, this shows that the discount factors in this framework are chosen in such a way as to satisfy the restrictions imposed by one bilateral country pair at the time.

To show this more clearly, Figure 4.3 compares the log levels of the discount factor for each country relative to each of the other two countries. For example, the first plot compares the time path of the log level discount factor for the USA when UK and Japan are used as partner countries, respectively. This time plot suggests that the discount factor for the USA is not uniquely determined, but clearly depends on the second country. Moreover, the differences between discount factors for the same country are the smallest for the USA and the largest for Japan, reflecting the relative importance of each country's excess return shocks on the log level of the discount factor.

Figure 4.4 presents scatterplots for the growth rates of the discount factors for each country when the other two countries are used as partners. The evidence in these scatterplots gives additional support to the findings from Figure 4.3. First, the measures for marginal utility growth (discount factor growth) for the same country are far from perfect.[25] Second, this imprecision in the measurement of discount factor growth increases with the "marginalization" of certain country's stock market shocks in the discount factor calculation. Hence, these measures are the least precise for Japan because it is the country with the lowest Sharpe ratio, and therefore, with the lowest discount factor loading (see Table 4.3). On the contrary, the imprecision is the lowest for the USA because this is the dominant country (highest Sharpe ratio and discount factor loading) in both country-pairs.

There is an intuitive interpretation of these findings as well. If an investor holds a portfolio of three risky assets with different risk-return profiles, then the asset that makes up the largest part of his total utility/well-being (highest Sharpe ratio) is the most important one for (the change in) his utility (represented by the stochastic discount factor). Following this argument, the contribution of the inferior asset (Japanese stock in this case) for investor's utility is very limited. Therefore, assets with relatively low Sharpe ratios represent residual assets for the investor. In turn, their contribution for his overall utility is quantified in a less precise manner.

Overall, the results suggest two main limitations of the bilateral SDF approach to international risk-sharing. First, the discount factors in the bilateral setting are not uniquely determined and show high sensitivity to the choice of particular partner country. Second, this sensitivity is especially important for countries with relatively low Sharpe ratios (on their stock markets), since their discount factors change substantially from one bilateral setting to another.

4.4 Trilateral Setting

In general, the discount factor for a certain country should be uniquely determined and incorporate all (direct) investment opportunities available to its residents (and therefore, should price all these assets). In order to investigate to what extent the results from section 4.3 depend on the specific, bilateral structure, we extend it into a three-country (trilateral) setting.[26] Therefore,

[25] Uniquely determined discount factors imply perfect relationships in all scatterplots, i.e. all points should lie along the 45 degrees line.

[26] All calculations for the trilateral setting can be found in the appendix.

Table 4.4: Real X-Rate and Discount Factor Volatility (Annualized)

Real X-Rate		Discount Factor	
e_1 (USA/UK)	11.58	USA	77.56
e_2 (USA/JAP)	12.47	UK	79.19
e_3 (JAP/UK)	12.05	JAP	76.05

Note: The table presents results for the components of the risk-sharing index in the trilateral setting. The first column gives figures for the the volatility of the real exchange rate, while the second column refers to the volatility of the stochastic discount factors over the time period August 1978-June 1998. Both volatilities (of the real exchange rate and the marginal utility growth) are measured as annualized standard deviations and are expressed in percentage terms (rounded to two decimal places). Real exchange rate e_1 is defined as the price of UK goods in terms of USA goods, i.e. the ratio of prices in the UK over prices in the USA ($e_1 = S^{\$/£}(P^{UK}/P^{USA})$). Similarly, e_2 is ratio of Japanese over USA prices and e_3 is ratio of UK over Japanese prices.

the discount factors calculated in this trilateral setting are *unique* for each country and *simultaneously* price all assets available to its residents (all risky assets in each of the three countries).[27]

4.4.1 Results from the Trilateral Setting

Table 4.4 presents figures for the real exchange rate and discount factor volatilities in the trilateral setting. Similar as in the bilateral case, marginal utility growth volatility is several times larger (about $70-80$ percent, measured by the discount factor volatility) than real exchange rate volatility (about 12 percent), suggesting that a lot of risk-sharing takes place among them.

We modify the risk-sharing index given by equation (4.13) in order to adapt it to our trilateral framework. Hence, we include all three countries in its calculation and calculate the trilateral risk-sharing index as follows:

$$RSI = 1 - \frac{\alpha\Sigma^{e_1 e_1} + (1-\alpha)\Sigma^{e_2 e_2}}{\mu^{d\prime}\Sigma_d^{-1}\mu^d + \alpha\mu^{f_1'}\Sigma_{f_1}^{-1}\mu^{f_1} + (1-\alpha)\mu^{f_2'}\Sigma_{f_2}^{-1}\mu^{f_2}} \tag{4.26}$$

For example, for the domestic country (USA), we include both real exchange rates (with respect to the UK and with respect to Japan) and all three discount factor volatilities. Moreover, we allow for differences between partner countries by assigning them specific weights α and $(1-\alpha)$, respectively. In this way, all foreign partner weights for a certain country must sum up to 1. The

[27] The extension to an n-country (n-assets) setting follows the same lines.

easiest way to think about this approach is as an "effective, trade-weighted" combination of foreign partners.

In fact, these weights should correspond to the relative importance of specific partner countries for international risk-sharing. Hence, there is no specific theoretical way to derive them.[28] Rather, in this study we allow the value for α to fluctuate anywhere between 0 and 1. Figure 4.5 shows results for the risk-sharing index for each country when different weights are assigned to its other two partners. In fact, the value for α, indicated on the horizontal axis, goes from one extreme (0) to the other (1) (where at each extreme only one of the partner countries matters for risk-sharing) and covers all possible intermediate cases.

For example, the line for the USA represents different values for the USA risk-sharing index going from $\alpha = 0$ (all risk-sharing is done with Japan) to $\alpha = 1$ (all risk-sharing takes place with the UK). The upward slope of this line with respect to α suggests that USA achieves a higher level of international risk-sharing when UK becomes the relatively more important partner. The similar logic applies to the calculations for the other two countries: the upward line for the UK indicates increasing risk-sharing levels when USA becomes relatively more important partner (compared to Japan), and the downward sloping line for Japan indicates decreasing risk-sharing levels when USA becomes relatively more important partner (compared to the UK).

We can derive two conclusions from this figure. First, though differences exist, the risk-sharing index does not vary a lot with respect to the specific combination of partner countries.[29] Second, irrespective of the relative importance of different partner countries, the risk-sharing index for each country-pair is higher than the corresponding index in the bilateral setting. This is the central result from our trilateral setting: measures of risk-sharing based on the stochastic discount factor approach are not very sensitive to the number of countries used in their calculation. If anything, then this trilateral framework suggests somewhat higher risk-sharing compared to the bilateral setting.

[28] For example, they can be calculated according to the share of trade or the portion of a country's assets portfolio invested in each country.

[29] The index fluctuates within the range 0.9865-0.9885.

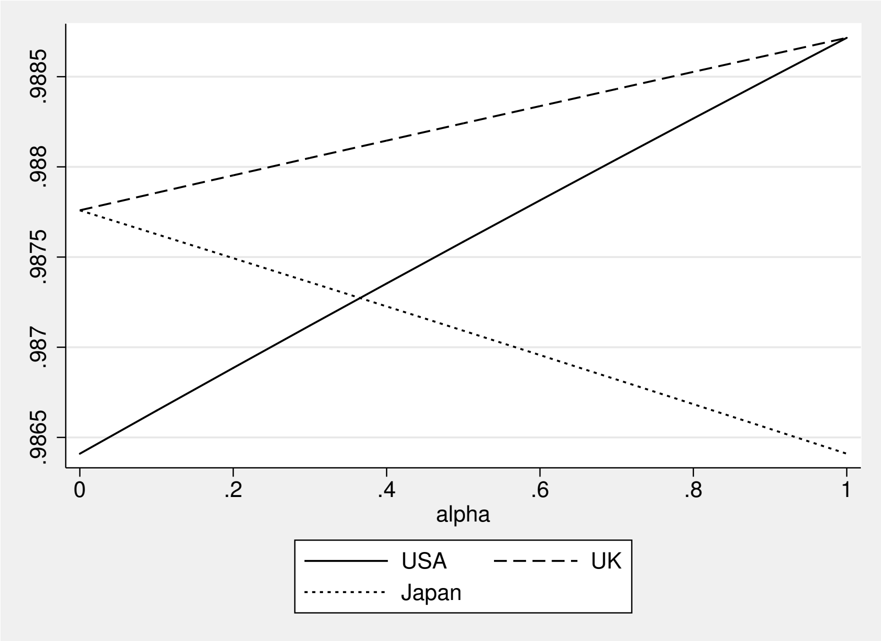


Figure 4.5: Risk-Sharing Index (Trilateral)

Note: The figure presents values for the risk-sharing index obtained from the trilateral setting given in equation (4.26): $RSI = 1 - \frac{\alpha \Sigma^{e_1 e_1} + (1-\alpha)\Sigma^{e_2 e_2}}{\mu^{d'} \Sigma_d^{-1} \mu^d + \alpha \mu^{f_1'} \Sigma_{f_1}^{-1} \mu^{f_1} + (1-\alpha) \mu^{f_2'} \Sigma_{f_2}^{-1} \mu^{f_2}}$. Each line refers to values of the risk-sharing index for one country when different weights (α) are assigned to its other two partners. For $\alpha = 0$ and $\alpha = 1$, the index measures risk-sharing between two countries. For $\alpha = 0$ the index refers to the following pairs: USA-Japan, UK-Japan, and Japan-UK, while for $\alpha = 1$ the index refers to the following pairs: USA-UK, UK-USA, and Japan-USA.

4.4.2 Visual Evidence

Figure 4.6 depicts the development of log discount factors through time. In this setting, all three discount factors are *simultaneously and uniquely* determined. As can be seen from the figure, their behavior closely resembles that for the bilateral country pairs. In fact, all three log discount factors move very closely together, the only difference being assigned to the fluctuations in the real exchange rates.

Finally, we complete our visual inspection with a 3-dimensional scatterplot of the discount factor growth rates given in Figure 4.7. In fact, this plot visualizes the joint correlation among the discount factor growth rates for all

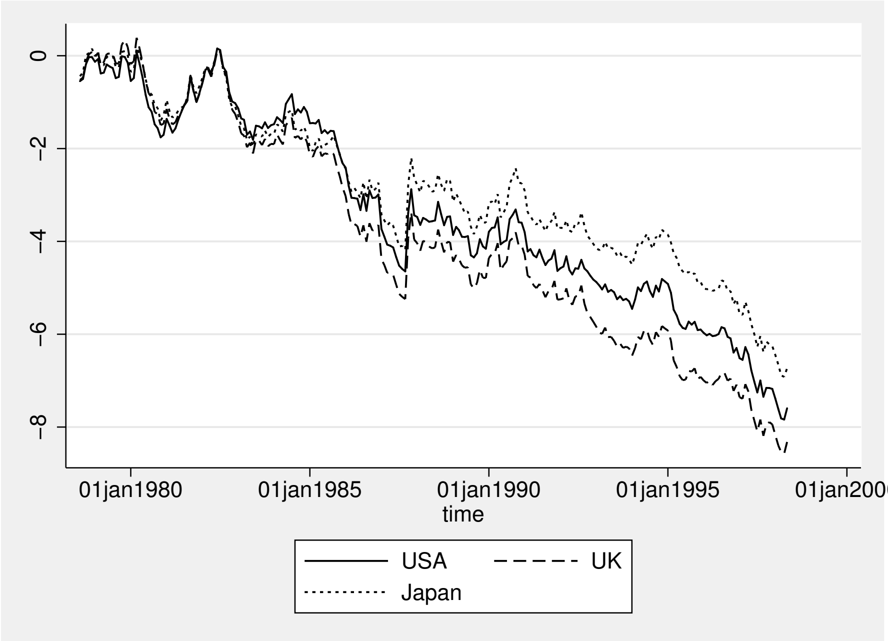


Figure 4.6: Log Level of Discount Factors

Note: The figure presents time lines of the log levels of the discount factors calculated in the trilateral setting. The log levels of the discount factors are calculated through accumulation of the changes in the log discount factors given in equation (4.20). All log levels are uniquely determined and price all assets in each of the three countries.

three countries. The figure shows that almost all points (observations) lie along the spatial diagonal, suggesting quasi-equalization of all three discount factor growth rates. Thus, the evidence from this 3-dimensional scatterplot just strengthens the conclusion that the stochastic discount factor approach implies somewhat higher international risk-sharing in the trilateral than it does in the bilateral setting.

4.5 Discussion: Limitations of the SDF Approach

In this section we discuss two main (possible) limitations of the SDF approach: the first refers to the internal incoherence of the bilateral approach, while the second refers to the discrepancy of its results with the macroeconomic evidence.

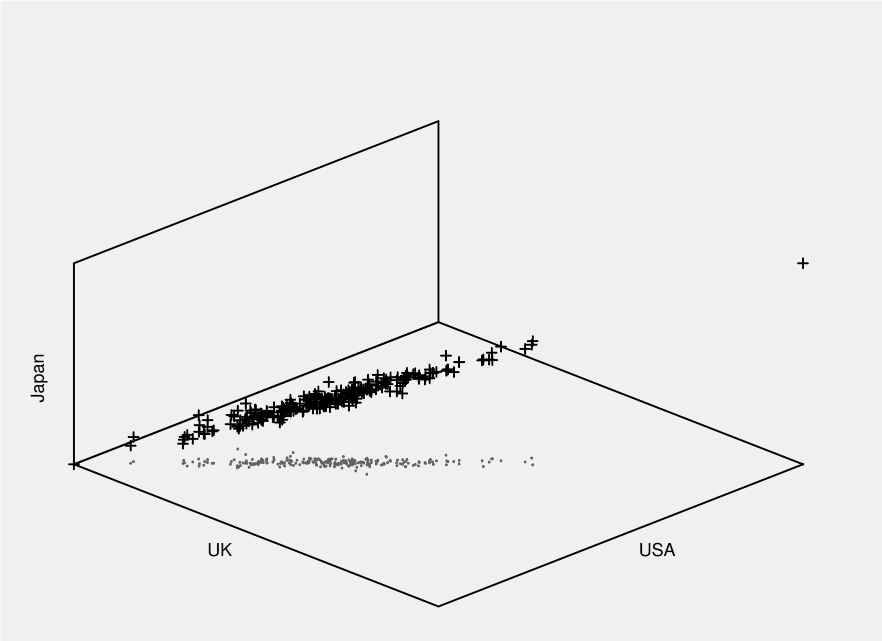


Figure 4.7: Correlation of Discount Factor Growth Rates (3-D)

Note: The figure presents a threedimensional scatterplot for growth rates of the discount factors calculated in the trilateral setting. The growth of discount factors is calculated according to equation (4.19).

4.5.1 Inherent Incoherence

The trilateral framework, presented in the previous section, tries to account for some of the (possible) criticisms about the inherent incoherence of the bilateral SDF approach identified in section 4.3.4. However, although it demonstrates that the (final) results for the international risk-sharing index are quite robust to the number of countries used in their calculation, the trilateral framework does not (completely) resolve the inherent incoherence found in the bilateral SDF model. In fact, it only (temporarily) fixes the problems from the bilateral setting, and therefore, faces the same type of (incoherence) criticisms with the addition of new countries.

Figure 4.6 shows that the discount factors calculated using the trilateral framework will be uniquely determined and price all assets in a three-country world. However, as soon as a fourth country enters this world, the trilateral framework will face the same problem(s) as the bilateral one.[30] In that case, the stochastic discount factors calculated from the trilateral setting will not be uniquely determined and will not price all assets simultaneously anymore. In fact, the addition of a fourth country brings two additional (independent) shocks into the system that cannot be (uniquely) priced by the discount factors computed from the trilateral setting.[31] Moreover, the time paths of the three discount factors in Figure 4.6 are very similar with the time paths of the discount factors for the bilateral pairs that include USA (the first and second panel in Figure 4.1). This observation suggests that the discount factors for all three countries crucially depend on the shock with the highest Sharpe ratio, i.e. the USA stock market shock again.[32] Hence, including a fourth country with a Sharpe ratio (for its stock market) even higher than the one for the USA might lead to dramatic changes in all discount factors from the trilateral setting. Following the same argument as for the bilateral setting in section 4.3.4, the discount factors computed from one trilateral setting (group of three countries) will (in general) be quite different from the discount factors computed from another trilateral setting in that case (another combination of three countries).[33] Therefore, the trilateral setting only shifts the problem with the internal incoherence of the SDF approach to international risk-sharing by one country ahead, but does not resolve it.

4.5.2 Reconciliation with Macroeconomic Evidence

Unambiguously, the results from the trilateral framework just strengthen the evidence about the discrepancy that exists between the measures of international risk-sharing derived from asset markets data following the stochastic

[30] The trilateral framework, elaborated on in section B.2, offers a simple extension to calculate risk-sharing among a group of several countries.

[31] Each country adds two additional (independent) shocks: one related to its stock market, and the other related to its foreign exchange market.

[32] Table B.1 in the appendix shows that all three discount factors load much more on the USA stock market shock than on the other shocks.

[33] There will be a total of four different three- country groups/combinations in this four-country world. Three of these groups will be strongly influenced by the highest Sharpe ratio country, and therefore, will be quite different from the last group/combination that excludes this country.

discount factor approach and those derived with macroeconomic data and specific utility function. One possible reason for these differences is the absence of complete capital markets. In fact, if asset markets account for only a small portion of total macroeconomic risks, then the low values for international risk-sharing implied by macroeconomic data can co-exist with the high risk-sharing measures presented here. However, these additional, non-marketable/non-insurable shocks not spanned by assets markets should be very large, negatively correlated across countries, and even more variable than the ones already observed in asset markets. In fact, Brandt et al. (2006) demonstrate that it is extremely difficult to justify the existence of such shocks.[34] Subsequently, shocks must be even larger and more variable to rationalize the results from the trilateral setting presented in this study. Therefore, it is very unlikely that the reconciliation between these two approaches to measuring international risk-sharing would go along these lines.

The arguments above suggest that equation (4.10) cannot hold if the two different approaches are to be reconciled.[35] In fact, by assuming that equation (4.10) holds at any point in time, i.e. that real exchange rates depreciate/appreciate exactly by the difference between domestic and foreign marginal utility growth rates, this approach implicitly "imposes" (almost) perfect risk-sharing. Compared to the macroeconomic literature, this condition is equivalent to the risk-sharing condition in the presence of non-traded goods proposed by Backus and Smith (1993): marginal utility growth (usually measured through consumption growth) can differ across countries as long as real exchange rates do not stay constant.[36] In that case, the Backus-Smith condition suggests that real exchange rates appreciate for countries that experience relatively higher marginal utility growth rates (relatively lower consumption

[34]Brandt et al. (2006) show that these additional, non-insurable shocks should be very volatile (adding 50-100 percent volatility in marginal utility growth per year) and poorly (or negatively) correlated in order to reconcile the risk-sharing figures from the SDF approach with those found in the macroeconomic studies.

[35]Equation (4.10) need not always hold in the presence of incomplete markets. However, there are (infinitely many) combinations of non-insurable shocks in the two countries, for which equation (4.10) still holds in an incomplete markets setting.

[36]For exposition of this risk-sharing condition see Backus and Smith (1993), Kollmann (1995), Ravn (2001), or Corsetti, Dedola, and Leduc (2007).

growth rates).

It is important to realize that the basic asset pricing equation (4.10) gives an equilibrium, no-arbitrage condition between three macroeconomic variables. Nonetheless, none of these variables refers to asset(s) that is continually traded on the asset markets.[37] Instead, all variables correspond to abstract concepts about aggregate macroeconomic behavior, which prevents direct empirical testing of this condition. Therefore, it might be interesting to test not only whether this condition holds as parity (as assumed here), but rather to see whether it has the correct sign (+). If this is not the case, then the reconciliation of the two approaches might be very closely related to the solution(s) of other puzzles in international macroeconomics and finance: the uncovered interest parity (UIP) anomaly and the Backus-Smith puzzle (consumption-real exchange rate correlation puzzle).

4.6 Concluding Remarks

In this study we present an extension of the stochastic discount factor approach to international risk-sharing. At the beginning, we present the theoretical framework that links the minimum-variance discount factors in two countries with the corresponding real exchange rate. We elaborate on the calculation of the discount factors, the construction of the risk-sharing index and the replication of the results for the bilateral setting given in Brandt et al. (2006). There are two possible criticisms about the inherent inconsistency of the bilateral approach. First, the discount factors are not uniquely determined in the bilateral framework and crucially depend on the partner country included in the calculation. Second, the deviations between the discount factors obtained in this way (the imprecision in the measurement of marginal utility growth) are larger for countries whose stock market excess return shocks are relatively less important. Both of these criticisms suggest that the (bilateral) SDF approach to international risk-sharing is very sensitive to the choice of

[37] Condition (4.10) was derived under the assumptions that there is free trade in all assets and there are no pure (zero initial investment) arbitrage opportunities. Therefore, the nature of the three macroeconomic variables used in condition (4.10) seriously questions both of these assumptions.

particular partner countries.

In order to account for some of these shortcomings of the bilateral framework, we propose an extension to a three-country (trilateral) setting. However, although the trilateral framework demonstrates that the (final) results for the international risk-sharing index are quite robust to the number of countries used in their calculation, it does not resolve the inherent incoherence found in the bilateral SDF model. In fact, as soon as a fourth country enters this world, the trilateral framework will face the same problem(s) as the bilateral one: the discount factors will not be uniquely determined and their behavior will crucially depend on the shock with the highest Sharpe ratio. Therefore, we conclude that the trilateral setting only shifts the problem with the internal incoherence of the SDF approach to international risk-sharing by one country ahead, but does not resolve it.

Finally, we give a note of caution on the interpretation of the results in this study. The stochastic discount factor approach to international risk-sharing is derived under the assumption that equation (4.10) always holds. Moreover, the replication of the results for the bilateral setting, but also the extension to a trilateral setting are performed retaining the assumption that equation (4.10) prices all assets at any point in time. However, if this is not the case, i.e. if the economies are far-away from what is implied by the first principles, then this approach cannot give valid measures of international risk-sharing in the first place.

Appendix B

B.1 Excess Returns in Bilateral Setting

This section presents formulae for excess returns in the bilateral framework. First, we present a general derivation for excess return formulae for each asset. Second, we derive vectors of expected excess returns for each country. A general distinction is made between formulae for domestic country (with superscript d) and foreign country (with superscript f) assets. USA is the domestic country in the first two country-pairs, and UK is the domestic country in the last country-pair.

B.1.1 Excess Return Processes for Domestic Investor

The investors in the domestic country face the following three types of excess return shocks: domestic stock, foreign bond, and foreign stock.

The excess returns on domestic stock are calculated difference between returns on the domestic stock market and the risk-free rate on domestic bond:

$$\frac{dS^d}{S^d} - \frac{dB^d}{B^d} = (\theta^d - r^d)dt + dz^d \tag{B.1}$$

The corresponding excess return on the foreign bond for the domestic investor is given as the difference between foreign bond return expressed in domestic currency and domestic bond return. Hence, although foreign bond is risk-free for the foreign investor, it is risky asset from the perspective of the domestic investor due to the currency risk it contains.

$$\frac{d(eB^f)}{eB^f} - \frac{dB^d}{B^d} = \frac{de}{e} + r^f dt - r^d dt = (\theta^e + r^f - r^d)dt + dz^e \tag{B.2}$$

Finally, excess returns on foreign stock for the domestic investor are calculated as the difference between returns on foreign stock and returns on foreign bonds when both are expressed in domestic currency units. Excess returns on the foreign stock for the domestic investor is calculated as follows:

$$\begin{aligned}
&\frac{d(eS^f)}{eS^f} - \frac{d(eB^f)}{eB^f} = \frac{dS^f}{S^f} + \frac{de}{e}\frac{dS^f}{S^f} - \frac{dB^f}{B^f} - \frac{de}{e}\frac{dB^f}{B^f} \\
&\quad = \Big(1 + \frac{de}{e}\Big)\Big(\frac{dS^f}{S^f} - \frac{dB^f}{B^f}\Big) \\
&\quad = (1 + \theta^e dt + dz^e)(\theta^f dt + dz^f - r^f dt) \\
&\quad = \theta^f dt + dz^f - r^f dt \\
&\quad + \theta^e dt\theta^f dt + \theta^e dt dz^f - \theta^e dt r^f dt + dz^e\theta^f dt + dz^e dz^f - dz^e r^f dt \\
&\quad = (\theta^f - r^f)dt + dz^e dz^f + dz^f \\
&\quad = (\theta^f - r^f + \Sigma^{ef})dt + dz^f
\end{aligned} \tag{B.3}$$

B.1.2 Excess Return Processes for Foreign Investor

The investors in the foreign country face the following three types of excess return shocks: domestic bond, domestic stock, and foreign stock.

The excess return on the domestic bond for the foreign investor is given as the difference between domestic bond return expressed in domestic currency and foreign bond return. Hence, although domestic bond is risk-free for the domestic investor, it is a risky asset from the perspective of the foreign investor due to the currency risk it contains. These excess returns are given as follows:

$$\begin{aligned}
&\frac{d\Big(\frac{B^d}{e}\Big)}{\Big(\frac{B^d}{e}\Big)} - \frac{dB^f}{B_f} = \Big(\frac{dB^d}{B^d} - \frac{de}{e} + \frac{de_1^2}{e_1^2} - \frac{de}{e}\frac{dB^d}{B^d}\Big) - \frac{dB^f}{B_f} \\
&\quad = r^d dt - \theta^e dt - dz^e + \Sigma^{ee} dt - \theta^d dt r^d dt - r^f dt \\
&\quad = (r^d - r^f - \theta^e + \Sigma^{ee})dt - dz^e \\
&\quad = -[(\theta^e + r^f - r^d - \Sigma^{ee})dt + dz^e]
\end{aligned} \tag{B.4}$$

The excess returns on domestic stock from the perspective of foreign investor are calculated as difference between domestic stock and domestic bond returns, both translated into foreign currency:

$$
\begin{aligned}
\frac{d\left(\frac{S^d}{e}\right)}{\frac{S^d}{e}} - \frac{d\left(\frac{B^d}{e}\right)}{\frac{B^d}{e}} &= \Big(\frac{dS^d}{S^d} - \frac{de}{e} + \frac{de_1^2}{e_1^2} - \frac{de}{e}\frac{dS^d}{S^d}\Big) \\
&- \Big(\frac{dB^d}{B^d} - \frac{de}{e} + \frac{de_1^2}{e_1^2} - \frac{de}{e}\frac{dB^d}{B^d}\Big) \\
&= \frac{dS^d}{S^d} - \frac{dB^d}{B^d} - \frac{de}{e}\Big(\frac{dS^d}{S^d} - \frac{dB^d}{B^d}\Big) \\
&= (1 - \frac{de}{e})\Big(\frac{dS^d}{S^d} - \frac{dB^d}{B^d}\Big) \\
&= (1 - \theta^e dt - dz^e)(\theta^d dt + dz^d - r^d dt) \\
&= \theta^d dt + dz^d - r^d dt - \theta^e dt \theta^d dt - \theta^e dt dz^d \\
&+ \theta^e dt r^d dt - dz^e \theta^d dt + dz^e dz^d - dz^e r^d dt \\
&= (\theta^d - r^d - \Sigma^{ed})dt + dz^d
\end{aligned}
\qquad \text{(B.5)}
$$

Finally, the excess returns that foreign investors get by investing on the foreign stock market are given as the difference between returns on foreign stock market and returns on foreign bond. Since the latter is a risk-free asset from the perspective of foreign investors. Hence, the foreign stock market excess returns are given by the following equation:

$$
\frac{dS^f}{S^f} - \frac{dB^f}{B^f} = (\theta^f - r^f)dt + dz^f \qquad \text{(B.6)}
$$

B.1.3 Expected Excess Returns

This section presents the expected values for the excess return processes calculated in the previous two sections. The term in front of the dt term refers to the expected values in the continuous-time formulation employed here. Therefore, domestic investor faces the following set of expected excess returns:

$$\mu^d = \begin{bmatrix} \theta^d - r^d \\ \theta^e + r^f - f^d \\ \theta^f - r^f + \Sigma^{ef} \end{bmatrix}$$

This vector stacks the expected values of the expected return processes given by equations (B.1) (domestic stock), (B.2) (foreign bond), and (B.6) (foreign stock).

The foreign investor faces a similar set of expected excess returns. The following vector stack the expected values of the expected return processes given by equations (B.4) (domestic bond), (B.5) (domestic stock), and (B.6) (foreign stock):

$$\mu^f = \begin{bmatrix} \theta^d - r^d - \Sigma^{ed} \\ -(\theta^e + r^f - r^d - \Sigma^{ee}) \\ \theta^f - r^f \end{bmatrix}$$

B.2 Calculations for the Trilateral Setting

This section presents calculations for the trilateral framework. The discount factors in this trilateral setting can be calculated according to equations (4.19) and (4.20):

$$\frac{d\Lambda^i}{\Lambda^i} = -r^i dt - \mu^{i\prime}\Sigma_i^{-1} dz_i, i = d, f_1, f_2$$

$$d\ln\Lambda = \frac{d\Lambda}{\Lambda} - \frac{1}{2}\frac{d\Lambda^2}{\Lambda^2} = -\Big(r + \frac{1}{2}\mu'\Sigma_i^{-1}\mu\Big)dt - \mu'\Sigma_i^{-1}dz_i$$

and their volatility according to equation (4.21):

$$\frac{1}{dt}\sigma^2(d\ln\Lambda^i) = \mu'\Sigma_i^{-1}\mu, i = d, f_1, f_2$$

where d refers to the domestic country, f_1 to the first foreign country, and f_2 to the second foreign country. In the calculations below, d stands for the USA, f_1 for the UK, and f_2 for Japan. In the trilateral setting, residents in each country are faced with five (instead of three) sources of uncertainty. Apart from shocks to domestic risky assets, they face two exchange rate shocks, and two foreign risky assets shocks. Thus, all these sources of uncertainty can be summarized in the following three vectors, each referring to residents of the corresponding country:

$$dz_d = \begin{bmatrix} dz^d \\ dz^{e_1} \\ dz^{e_2} \\ dz^{f_1} \\ dz^{f_2} \end{bmatrix}$$

$$dz_{f_1} = \begin{bmatrix} dz^d \\ dz^{e_1} \\ dz^{e_3} \\ dz^{f_1} \\ dz^{f_2} \end{bmatrix}$$

$$dz_{f_2} = \begin{bmatrix} dz^d \\ dz^{e_3} \\ dz^{e_2} \\ dz^{f_1} \\ dz^{f_2} \end{bmatrix}$$

with the following set of three variance-covariance matrices:

$$\Sigma_d = \frac{1}{dt}E(dz_d dz_d') = \begin{bmatrix} \Sigma^{dd'} & \Sigma^{de_1} & \Sigma^{de_2} & \Sigma^{df_1'} & \Sigma^{df_2'} \\ \Sigma^{e_1 d'} & \Sigma^{e_1' e_1} & \Sigma^{e_1' e_2} & \Sigma^{e_1 f_1'} & \Sigma^{e_1 f_2'} \\ \Sigma^{e_2 d'} & \Sigma^{e_2' e_1} & \Sigma^{e_2' e_2} & \Sigma^{e_2 f_1'} & \Sigma^{e_2 f_2'} \\ \Sigma^{f_1' d} & \Sigma^{f_1' e_1} & \Sigma^{f_1' e_2} & \Sigma^{f_1' f_1} & \Sigma^{f_1' f_2} \\ \Sigma^{f_2' d} & \Sigma^{f_2' e_1} & \Sigma^{f_2' e_2} & \Sigma^{f_2' f_1} & \Sigma^{f_2' f_2} \end{bmatrix}$$

$$\Sigma_{f_1} = \frac{1}{dt}E(dz_{f_1} dz_{f_1}') = \begin{bmatrix} \Sigma^{dd'} & \Sigma^{de_1} & \Sigma^{de_3} & \Sigma^{df_1'} & \Sigma^{df_2'} \\ \Sigma^{e_1 d'} & \Sigma^{e_1' e_1} & \Sigma^{e_1' e_3} & \Sigma^{e_1 f_1'} & \Sigma^{e_1 f_2'} \\ \Sigma^{e_3 d'} & \Sigma^{e_3' e_1} & \Sigma^{e_3' e_3} & \Sigma^{e_3 f_1'} & \Sigma^{e_3 f_2'} \\ \Sigma^{f_1' d} & \Sigma^{f_1' e_1} & \Sigma^{f_1' e_3} & \Sigma^{f_1' f_1} & \Sigma^{f_1' f_2} \\ \Sigma^{f_2' d} & \Sigma^{f_2' e_1} & \Sigma^{f_2' e_3} & \Sigma^{f_2' f_1} & \Sigma^{f_2' f_2} \end{bmatrix}$$

$$\Sigma_{f_2} = \frac{1}{dt}E(dz_{f_2} dz_{f_2}') = \begin{bmatrix} \Sigma^{dd'} & \Sigma^{de_3} & \Sigma^{de_2} & \Sigma^{df_1'} & \Sigma^{df_2'} \\ \Sigma^{e_3 d'} & \Sigma^{e_3' e_3} & \Sigma^{e_3' e_2} & \Sigma^{e_3 f_1'} & \Sigma^{e_3 f_2'} \\ \Sigma^{e_2 d'} & \Sigma^{e_2' e_3} & \Sigma^{e_2' e_2} & \Sigma^{e_2 f_1'} & \Sigma^{e_2 f_2'} \\ \Sigma^{f_1' d} & \Sigma^{f_1' e_3} & \Sigma^{f_1' e_2} & \Sigma^{f_1' f_1} & \Sigma^{f_1' f_2} \\ \Sigma^{f_2' d} & \Sigma^{f_2' e_3} & \Sigma^{f_2' e_2} & \Sigma^{f_2' f_1} & \Sigma^{f_2' f_2} \end{bmatrix}$$

Moreover, we must impose an additional restriction in the calculation. Namely, we have to exclude the possibilities for triangular (cross-currency) arbitrage. In particular, if the exchange rate returns are given by:

$$\frac{de_1}{e_1} = \theta_1^e dt + dz_1^e, \frac{de_2}{e_2} = \theta_2^e dt + dz_2^e, \frac{de_3}{e_3} = \theta_3^e dt + dz_3^e \tag{B.7}$$

then the following cross-currency condition must hold:

$$\theta_3^e dt + dz_3^e = \theta_2^e dt + dz_2^e + \theta_1^e dt + dz_1^e \tag{B.8}$$

The excess return vectors can be related using the restrictions imposed by the cross-currency condition (B.8) (no triangular arbitrage possibilities). For example, the excess returns for a domestic resident can be related with the excess returns for a resident in the first foreign country (f_1) as follows:[1]

$$\mu^{f_1} = A\mu^d \tag{B.9}$$

where the matrix A is defined as:

$$A = \begin{bmatrix} 1 & 0 & 0 & 0 & 0 \\ 0 & -1 & -1 & 0 & 0 \\ 0 & 0 & -1 & 0 & 0 \\ 0 & 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 0 & 1 \end{bmatrix} \tag{B.10}$$

Equation (B.9) shows that residents in both countries face the same expected excess returns on all three stock markets, while their foreign exchange excess returns form a linear combination. In turn, the variance covariance matrix with shocks facing the residents in the first foreign country is given by:

$$\Sigma_{f_1} = A\Sigma_d A' \tag{B.11}$$

and its inverse:

$$\Sigma_{f_1}^{-1} = (A\Sigma_d A')^{-1} = (A')^{-1}\Sigma_d^{-1}A^{-1} \tag{B.12}$$

Therefore the domestic and first foreign (f_1) discount factor loadings will be related as follows:

$$\mu^{f_1}\Sigma_{f_1}^{-1} = \mu^d A'(A')^{-1}\Sigma_d^{-1}A^{-1} = \mu^d\Sigma_d^{-1}A^{-1} \tag{B.13}$$

Equation (B.13) indicates that the only difference between domestic and foreign discount factors is given by A^{-1}. It means that all discount factors load equally on all three stock market shocks, while their foreign exchange loadings differ by a linear combination of the exchange rate shocks.

[1] For reasons of symmetry we use directly the discrete-time equivalents of the continuous-time formulae, just as implemented in the calculations. Thus, we disregard the continuous-time terms in the excess return vectors.

B.2.1 Excess Returns in Trilateral Setting

This section presents formulae for excess returns in the trilateral framework. Similar as in the bilateral case, we present a derivation for (expected) excess returns on each asset. A general distinction is made between formulae for domestic country (with superscript d) and two foreign countries (with superscript f_1 and f_2). USA is the domestic country, UK refers to foreign country f_1 and Japan refers to foreign country f_2.

Domestic - USA

A USA-based resident gets the following excess returns on the domestic stock market:

$$\frac{dS^d}{S^d} - \frac{dB^d}{B^d} = (\theta^d - r^d)dt + dz^d \tag{B.14}$$

similarly, he gets the following excess returns on the foreign bond in country f_1 (UK):

$$\begin{aligned}\frac{d\left(\frac{B^d}{e_1}\right)}{\left(\frac{B^d}{e_1}\right)} - \frac{dB^{f_1}}{B_{f_1}} &= \Big(\frac{dB^d}{B^d} - \frac{de_1}{e_1} + \frac{de_1^2}{e_1^2} - \frac{de_1}{e_1}\frac{dB^d}{B^d}\Big) - \frac{dB^{f_1}}{B_{f_1}}\\ &= r^d dt - \theta^{e_1}dt - dz^{e_1} + \Sigma^{e_1e_1}dt - \theta^d dt r^d dt - r^{f_1}dt\\ &= (r^d - r^{f_1} - \theta^{e_1} + \Sigma^{e_1e_1})dt - dz^{e_1}\\ &= -[(\theta^{e_1} + r^{f_1} - r^d - \Sigma^{e_1e_1})dt + dz^{e_1}]\end{aligned} \tag{B.15}$$

on the foreign bond in Japan (country f_2):

$$\frac{d(e_2B^{f_2})}{e_2B^{f_2}} - \frac{dB^d}{B^d} = \frac{de_2}{e_2} + r^{f_2}dt - r^d dt = (\theta^{e_2} + r^{f_2} - r^d)dt + dz^{e_2} \tag{B.16}$$

on the stock market in the UK (country f_1):

$$\begin{aligned}&\frac{d(e_1S^{f_1})}{e_1S^{f_1}} - \frac{d(e_1B^{f_1})}{e_1B^{f_1}} = \frac{dS^{f_1}}{S^{f_1}} + \frac{de_1}{e_1}\frac{dS^{f_1}}{S^{f_1}} - \frac{dB^{f_1}}{B^{f_1}} - \frac{de_1}{e_1}\frac{dB^{f_1}}{B^{f_1}}\\ &\quad = \Big(1 + \frac{de_1}{e_1}\Big)\Big(\frac{dS^{f_1}}{S^{f_1}} - \frac{dB^{f_1}}{B^{f_1}}\Big)\\ &\quad = (1 + \theta^{e_1}dt + dz^{e_1})(\theta^{f_1}dt + dz^{f_1} - r^{f_1}dt)\end{aligned}$$

$$
\begin{aligned}
&= \theta^{f_1}dt + dz^{f_1} - r^{f_1}dt \\
&+ \theta^{e_1}dt\theta^{f_1}dt + \theta^{e_1}dtdz^{f_1} - \theta^{e_1}dtr^{f_1}dt + dz^{e_1}\theta^{f_1}dt + dz^{e_1}dz^{f_1} - dz^{e_1}r^{f_1}dt \\
&= (\theta^{f_1} - r^{f_1})dt + dz^{e_1}dz^{f_1} + dz^{f_1} \\
&= (\theta^{f_1} - r^{f_1} + \Sigma^{e_1 f_1})dt + dz^{f_1} \qquad \text{(B.17)}
\end{aligned}
$$

and on the stock market in Japan (country f_2):

$$
\begin{aligned}
&\frac{d(e_2S^{f_2})}{e_2S^{f_2}} - \frac{d(e_2B^{f_2})}{e_2B^{f_2}} = \frac{dS^{f_2}}{S^{f_2}} + \frac{de_2}{e_2}\frac{dS^{f_2}}{S^{f_2}} - \frac{dB^{f_2}}{B^{f_2}} - \frac{de_2}{e_2}\frac{dB^{f_2}}{B^{f_2}} \\
&= \Big(1 + \frac{de_2}{e_2}\Big)\Big(\frac{dS^{f_2}}{S^{f_2}} - \frac{dB^{f_2}}{B^{f_2}}\Big) \\
&= (1 + \theta^{e_2}dt + dz^{e_2})(\theta^{f_2}dt + dz^{f_2} - r^{f_2}dt) \\
&= \theta^{f_2}dt + dz^{f_2} - r^{f_2}dt \\
&+ \theta^{e_2}dt\theta^{f_2}dt + \theta^{e_2}dtdz^{f_2} - \theta^{e_2}dtr^{f_2}dt + dz^{e_2}\theta^{f_2}dt + dz^{e_2}dz^{f_2} - dz^{e_2}r^{f_2}dt \\
&= (\theta^{f_2} - r^{f_2})dt + dz^{e_2}dz^{f_2} + dz^{f_2} \\
&= (\theta^{f_2} - r^{f_2} + \Sigma^{e_2 f_2})dt + dz^{f_2} \qquad \text{(B.18)}
\end{aligned}
$$

Foreign 1 - UK

UK-based investor gets the following excess return on the USA (domestic) stock market:

$$
\begin{aligned}
&\frac{d\Big(\frac{S^d}{e_1}\Big)}{\frac{S^d}{e_1}} - \frac{d\Big(\frac{B^d}{e_1}\Big)}{\frac{B^d}{e_1}} = \Big(\frac{dS^d}{S^d} - \frac{de_1}{e_1} + \frac{de_1^2}{e_1^2} - \frac{de_1}{e_1}\frac{dS^d}{S^d}\Big) \\
&- \Big(\frac{dB^d}{B^d} - \frac{de_1}{e_1} + \frac{de_1^2}{e_1^2} - \frac{de_1}{e_1}\frac{dB^d}{B^d}\Big) \\
&= \frac{dS^d}{S^d} - \frac{dB^d}{B^d} - \frac{de_1}{e_1}\Big(\frac{dS^d}{S^d} - \frac{dB^d}{B^d}\Big) \\
&= (1 - \frac{de_1}{e_1})\Big(\frac{dS^d}{S^d} - \frac{dB^d}{B^d}\Big) \\
&= (1 - \theta^{e_1}dt - dz^{e_1})(\theta^d dt + dz^d - r^d dt) \\
&= \theta^d dt + dz^d - r^d dt - \theta^{e_1}dt\theta^d dt - \theta^{e_1}dtdz^d + \\
&\theta^{e_1}dtr^d dt - dz^{e_1}\theta^d dt + dz^{e_1}dz^d - dz^{e_1}r^d dt \\
&= (\theta^d - r^d - \Sigma^{e_1 d})dt + dz^d \qquad \text{(B.19)}
\end{aligned}
$$

and the following excess return on the USA (domestic) bond:

$$\frac{d(e_1 B^{f_1})}{e_1 B^{f_1}} - \frac{dB^d}{B^d} = \frac{de_1}{e_1} + r^{f_1}dt - r^d dt = (\theta^{e_1} + r^{f_1} - r^d)dt + dz^{e_1} \quad \text{(B.20)}$$

while investment in the UK (foreign f_1) stock market brings him the following excess return:

$$\begin{aligned}
\frac{d\left(\frac{B^{f_2}}{e_3}\right)}{\left(\frac{B^{f_2}}{e_3}\right)} - \frac{dB^{f_1}}{B_{f_1}} &= \Big(\frac{dB^{f_2}}{B^{f_2}} - \frac{de_3}{e_3} + \frac{de_3^2}{e_3^2} - \frac{de_3}{e_3}\frac{dB^{f_2}}{B^{f_2}}\Big) - \frac{dB^{f_1}}{B_{f_1}} \\
&= r^{f_2}dt - \theta^{e_3}dt - dz^{e_3} + \Sigma^{e_3 e_3}dt - \theta^{f_2}dt r^{f_2}dt - r^{f_1}dt \\
&= (r^{f_2} - r^{f_1} - \theta^{e_3} + \Sigma^{e_3 e_3})dt - dz^{e_3} \\
&= -[(\theta^{e_3} + r^{f_1} - r^{f_2} - \Sigma^{e_3 e_3})dt + dz^{e_3}] \qquad \text{(B.21)}
\end{aligned}$$

similar calculation can be made for Japanese bonds:

$$\frac{dS^{f_1}}{S^{f_1}} - \frac{dB^{f_1}}{B^{f_1}} = (\theta^{f_1} - r^{f_1})dt + dz^{f_1} \quad \text{(B.22)}$$

and investment on the Japanese stock market:

$$\begin{aligned}
\frac{d\left(\frac{S^{f_2}}{e_3}\right)}{\frac{S^{f_2}}{e_3}} - \frac{d\left(\frac{B^{f_2}}{e_3}\right)}{\frac{B^{f_2}}{e_3}} &= \Big(\frac{dS^{f_2}}{S^{f_2}} - \frac{de_3}{e_3} + \frac{de_3^2}{e_3^2} - \frac{de_3}{e_3}\frac{dS^{f_2}}{S^{f_2}}\Big) \\
&\quad - \Big(\frac{dB^{f_2}}{B^{f_2}} - \frac{de_3}{e_3} + \frac{de_3^2}{e_3^2} - \frac{de_3}{e_3}\frac{dB^{f_2}}{B^{f_2}}\Big) \\
&= \frac{dS^{f_2}}{S^{f_2}} - \frac{dB^{f_2}}{B^{f_2}} - \frac{de_3}{e_3}\Big(\frac{dS^{f_2}}{S^{f_2}} - \frac{dB^{f_2}}{B^{f_2}}\Big) \\
&= (1 - \frac{de_3}{e_3})\Big(\frac{dS^{f_2}}{S^{f_2}} - \frac{dB^{f_2}}{B^{f_2}}\Big) \\
&= (1 - \theta^{e_3}dt - dz^{e_3})(\theta^{f_2}dt + dz^{f_2} - r^{f_2}dt) \\
&= \theta^{f_2}dt + dz^{f_2} - r^{f_2}dt - \theta^{e_3}dt\theta^{f_2}dt - \theta^{e_3}dtdz^{f_2} + \theta^{e_3}dtr^{f_2}dt \\
&\quad - dz^{e_3}\theta^{f_2}dt + dz^{e_3}dz^{f_2} - dz^{e_3}r^{f_2}dt \\
&= (\theta^{f_2} - r^{f_2} - \Sigma^{e_3 f_2})dt + dz^{f_2} \qquad \text{(B.23)}
\end{aligned}$$

Foreign 2 - Japan

Japan-based investor gets the following excess return on the USA stock market:

$$
\begin{aligned}
\frac{d\Big(\frac{S^d}{e_2}\Big)}{\frac{S^d}{e_2}} - \frac{d\Big(\frac{B^d}{e_2}\Big)}{\frac{B^d}{e_2}} &= \Big(\frac{dS^d}{S^d} - \frac{de_2}{e_2} + \frac{de_2^2}{e_2^2} - \frac{de_2}{e_2}\frac{dS^d}{S^d}\Big) \\
&- \Big(\frac{dB^d}{B^d} - \frac{de_2}{e_2} + \frac{de_2^2}{e_2^2} - \frac{de_2}{e_2}\frac{dB^d}{B^d}\Big) \\
&= \frac{dS^d}{S^d} - \frac{dB^d}{B^d} - \frac{de_2}{e_2}\Big(\frac{dS^d}{S^d} - \frac{dB^d}{B^d}\Big) \\
&= (1 - \frac{de_2}{e_2})\Big(\frac{dS^d}{S^d} - \frac{dB^d}{B^d}\Big) \\
&= (1 - \theta^{e_2}dt - dz^{e_2})(\theta^d dt + dz^d - r^d dt) \\
&= \theta^d dt + dz^d - r^d dt - \theta^{e_2}dt\theta^d dt - \theta^{e_2}dtdz^d \\
&+ \theta^{e_2}dtr^d dt - dz^{e_2}\theta^d dt + dz^{e_2}dz^d - dz^{e_2}r^d dt \\
&= (\theta^d - r^d - \Sigma^{e_2 d})dt + dz^d
\end{aligned} \tag{B.24}
$$

on the USA bond:

$$
\begin{aligned}
\frac{d\Big(\frac{B^d}{e_2}\Big)}{\Big(\frac{B^d}{e_2}\Big)} - \frac{dB^{f_2}}{B_{f_2}} &= \Big(\frac{dB^d}{B^d} - \frac{de_2}{e_2} + \frac{de_2^2}{e_2^2} - \frac{de_2}{e_2}\frac{dB^d}{B^d}\Big) - \frac{dB^{f_2}}{B_{f_2}} \\
&= r^d dt - \theta^{e_2}dt - dz^{e_2} + \Sigma^{e_2 e_2}dt - \theta^d dtr^d dt - r^{f_2}dt \\
&= (r^d - r^{f_2} - \theta^{e_2} + \Sigma^{e_2 e_2})dt - dz^{e_2} \\
&= -[(\theta^{e_2} + r^{f_2} - r^d - \Sigma^{e_2 e_2})dt + dz^{e_2}]
\end{aligned} \tag{B.25}
$$

on UK (country f_1) bond:

$$
\frac{d(e_3 B^{f_1})}{e_3 B^{f_1}} - \frac{dB^{f_2}}{B^{f_2}} = \frac{de_3}{e_3} + r^{f_1}dt - r^{f_2}dt = (\theta^{e_3} + r^{f_1} - f^{f_2})dt + dz^{e_3} \tag{B.26}
$$

on the UK stock market:

$$
\begin{aligned}
\frac{d(e_3 S^{f_1})}{e_3 S^{f_1}} - \frac{d(e_3 B^{f_1})}{e_3 B^{f_1}} &= \frac{dS^{f_1}}{S^{f_1}} + \frac{de_3}{e_3}\frac{dS^{f_1}}{S^{f_1}} - \frac{dB^{f_1}}{B^{f_1}} - \frac{de_3}{e_3}\frac{dB^{f_1}}{B^{f_1}} \\
&= \Big(1 + \frac{de_3}{e_3}\Big)\Big(\frac{dS^{f_1}}{S^{f_1}} - \frac{dB^{f_1}}{B^{f_1}}\Big)
\end{aligned}
$$

$$
\begin{aligned}
&= (1 + \theta^{e_3} dt + dz^{e_3})(\theta^{f_1} dt + dz^{f_1} - r^{f_1} dt) \\
&= \theta^{f_1} dt + dz^{f_1} - r^{f_1} dt \\
&+ \theta^{e_3} dt \theta^{f_1} dt + \theta^{e_3} dt dz^{f_1} - \theta^{e_3} dt r^{f_1} dt + dz^{e_3} \theta^{f_1} dt + dz^{e_3} dz^{f_1} - dz^{e_3} r^{f_1} dt \\
&= (\theta^{f_1} - r^{f_1}) dt + dz^{e_3} dz^{f_1} + dz^{f_1} \\
&= (\theta^{f_1} - r^{f_1} + \Sigma^{e_3 f_1}) dt + dz^{f_1} \qquad \text{(B.27)}
\end{aligned}
$$

and on the Japanese stock market:

$$
\frac{dS^{f_2}}{S^{f_2}} - \frac{dB^{f_2}}{B^{f_2}} = (\theta^{f_2} - r^{f_2}) dt + dz^{f_2} \qquad \text{(B.28)}
$$

B.2.2 Expected Excess Returns

The set of expected excess returns for the USA-based investor is given in the following vector:

$$\mu^d = \begin{bmatrix} \theta^d - r^d \\ -(\theta^{e_1} + r^{f_1} - r^d - \Sigma^{e_1 e_1}) \\ \theta^{e_2} + r^{f_2} - r^d \\ \theta^{f_1} - r^{f_1} + \Sigma^{e_1 f_1} \\ \theta^{f_2} - r^{f_2} + \Sigma^{e_2 f_2} \end{bmatrix}$$

similarly for the set of expected excess returns facing the UK-based investor:

$$\mu^{f_1} = \begin{bmatrix} \theta^d - r^d - \Sigma^{e_1 d} \\ \theta^{e_1} + r^{f_1} - r^d \\ -(\theta^{e_3} + r^{f_1} - r^{f_2} - \Sigma^{e_3 e_3}) \\ \theta^{f_1} - r^{f_1} \\ \theta^{f_2} - r^{f_2} - \Sigma^{e_3 f_2} \end{bmatrix}$$

and for the Japan-based investor:

$$\mu^{f_2} = \begin{bmatrix} \theta^d - r^d - \Sigma^{e_2 d} \\ -(\theta^{e_2} + r^{f_2} - r^d - \Sigma^{e_2 e_2}) \\ \theta^{e_3} + r^{f_1} - f^{f_2} \\ \theta^{f_1} - r^{f_1} + \Sigma^{e_3 f_1} \\ \theta^{f_2} - r^{f_2} \end{bmatrix}$$

The interpretation of these excess returns is analogous to the one given for the bilateral setting. In fact, the main difference is that residents can invest in two (instead of one) foreign risk-free bonds and three (instead of two) stock markets.

Table B.1: Discount Factor Loadings (Trilateral)

		USA	UK	Japan
dz^d	(dz^{USA})	4.07	4.07	4.07
dz^{e_1}	$(dz^{USA/UK})$	-0.23	1.74	
dz^{e_2}	$(dz^{USA/JAP})$	1.51		0.23
dz^{e_3}	$(dz^{JAP/UK})$		-1.51	-1.74
dz^{f_1}	(dz^{UK})	1.86	1.86	1.86
dz^{f_2}	(dz^{JAP})	-0.45	-0.45	-0.45

Note: The table presents figures for the discount factor loadings in the trilateral setting. The loadings for the discount factor in country i are given by $\mu^{i\prime}\Sigma_i^{-1}$. There are three stock market shocks and three real exchange rate shocks in this trilateral framework. The row marked $dz^d(dz^{USA})$ contains figures for discount factor loadings on the USA stock market shocks, row $dz^{f_1}(dz^{UK})$ refers to discount factor loadings on the UK stock market shocks, and row $dz^{f_2}(dz^{JAP})$ refers to discount factor loadings on the Japanese stock market shock. Rows marked dz^{e_i} contain figures for discount factor loadings on shocks for real exchange rate i. $dz^{e_1}(dz^{USA/UK})$ is defined as the relative price of UK in terms of USA goods, i.e. as the ratio of UK price level of USA price level. Similar definitions apply to $dz^{e_2}(dz^{USA/JAP})$ and $dz^{e_3}(dz^{JAP/UK})$.

B.2.3 Discount Factor Loadings

The evolution of the stochastic discount factors in the trilateral framework depends on five excess return shocks: three associated with the stock markets in each country plus two associated with the exchange rates. Table B.1 presents the discount factor loadings on these five shocks for each of the three countries.

Several findings in this table deserve attention. First, in line with the results for the bilateral setting and equation (B.13), all discount factors load equally on the stock market excess return shocks in each country. Second, these loadings differ across stock markets, being the the strongest for the USA, and the weakest for Japan. In fact, the magnitude and the relative importance of these loadings on the stock markets in the trilateral setting closely resemble those for the two bilateral pairs in section 4.3. Third, as pointed out in equations (B.9) and (B.13), each exchange rate loading forms a linear combination of the other two. For example, condition (B.13) and the definition of matrix A given in (B.10) imply the following relation between the loadings on the exchange rate excess return shocks for the domestic (USA) and the first foreign country (UK): $dz_{f_1}^{e_2} = dz_d^{e_2} - dz_d^{e_1}$. The values in Table B.1

confirm this linear relationship: $(1.74 = 1.51 - (-0.23))$. Similar conclusions apply to the other exchange rate shock combinations given in the second, the third, and the fourth row of Table B.1.

B.2.4 Pairwise Comparisons of SDF Growth Rates

Figure B.1 plots the discount factor growth rates for all three country pairs (bilaterally). This figure is almost identical to Figure 4.2, which depicted the correlation of discount factor growth rates in the bilateral setting. As in the previous case, most observations lie on or very close to the 45 degrees lines, suggesting that marginal utility growth rates are almost equalized for each bilateral country pair. This is exactly what the perfect risk-sharing condition implies.

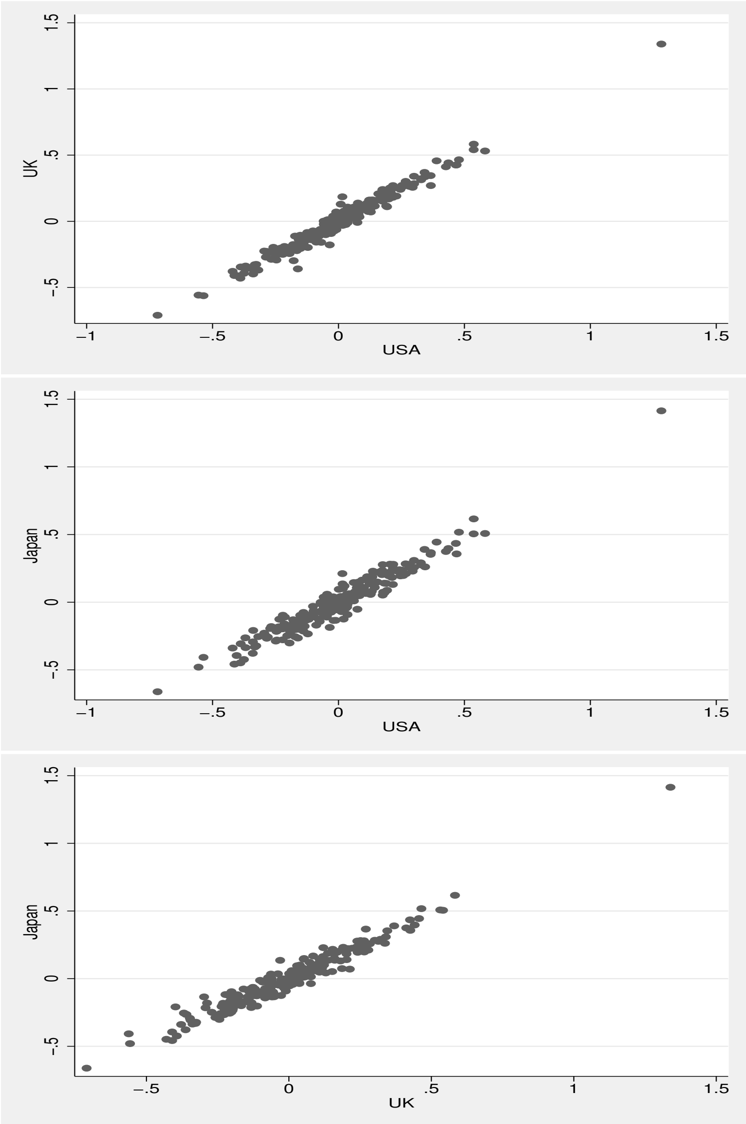


Figure B.1: Discount Factor Growth Rates (Trilateral)

Note: The figure presents scatterplots for growth rates of the discount factors calculated in the trilateral setting. Each plot refers to one of the three country-pair combinations. The growth of discount factors is calculated according to equation (4.19).

Chapter 5

Stochastic Discount Factor Approach to International Risk-Sharing: Evidence from Fixed Exchange Rate Episodes

5.1 Introduction

The stochastic discount factor model suggests that countries share a very high portion of all macroeconomic risks they face. Starting with the underlying assumption that real exchange rate changes equal the difference between domestic and foreign marginal utility growth (or discount factors), Brandt, Cochrane, and Santa-Clara (2006) argue that only a small portion of total risk is left unshared across countries. In fact, they calculate domestic and foreign marginal utility growth rates through stochastic discount factors derived from asset markets data (using the excess returns of the stock market index above the risk-free rate) and compare the volatility of these stochastic discount factors with the volatility of the real exchange rate. The main finding in their study is that the real exchange-rates (difference between marginal utility growth rates) are much less volatile than what the stochastic discount factors (proxies for marginal utility growth) of the corresponding countries would imply. Therefore, they conclude that marginal utility growth rates must be very highly correlated across countries, i.e. a large portion of macroeconomic risk is shared internationally.

The results from the approach employed in this study crucially depend on the behavior of the real exchange rate, which comprises two components: the ratio of price levels in the two countries, and the nominal exchange rate. Therefore, fluctuations in the real exchange rate can come from two sources: the behavior of the inflation rate differential, and the behavior of the nominal exchange rate changes. Hence, it might be interesting to see whether by fixing the (bilateral) nominal exchange rate, two countries "*automatically (or mechanically)*" achieve perfect levels of international risk-sharing. Does the stochastic discount factor approach imply perfect risk-sharing between countries when their nominal exchange rate is fixed?

We present some new evidence about the stochastic discount factor approach to measuring international risk-sharing when applied to episodes of fixed/rigid nominal exchange rates. We calculate risk-sharing indices for two such episodes and arrive at two main results: first, when applied to the period after the introduction of the Euro, this asset markets-based approach implies almost perfect risk-sharing for each Euro-zone country vis-*à*-vis the Euro-zone as a whole. This comes from the fact that the major part of fluctuations in the real exchange-rate (i.e. the difference in marginal utility growth rates across countries according to the asset pricing model) is eliminated after the introduction of the common currency. Second, when applied to emerging markets with fixed nominal exchange rates against the US dollar in the period 1993-2005, the approach again implies almost perfect risk-sharing for each country against the USA. In fact, these countries achieved much more risk-sharing vis-a-vis the US compared to developed countries like the UK or Japan.

Two main conclusions can be drawn from our findings: first, the risk-sharing measures crucially depend on the behavior of the nominal exchange rate, implying almost perfect risk-sharing among countries with fixed/rigid nominal exchange rates. Second, a counterintuitive ranking of risk-sharing levels under different nominal exchange rate regimes suggests a limited use of this approach for cross-country risk-sharing comparisons.

The rest of this chapter is organized as follows. Section 5.2 investigates the importance of fixed nominal exchange rate regimes: section 5.2.1 presents evidence for the Eurozone, while section 5.2.2 for several emerging economies. We discuss the relevance of our findings in section 5.3. Section 5.4 concludes the chapter.

5.2 Fixed Nominal Exchange Rates

The international risk-sharing index represents a comparison between real exchange rate volatility and the sum of the discount factor volatilities in the corresponding countries. The idea behind this set-up is that real exchange rates fluctuate by the difference between marginal utility growth rates in the corresponding countries, and thereby reflect hurdles for complete cross-country risk-sharing. Therefore, the results from the approach employed in this study crucially depend on the behavior of the real exchange rate. In turn, the real exchange rate is composed of two components: the ratio of price levels in the two countries, and the nominal exchange rate. Therefore, fluctuations in the real exchange rate can come from two sources: the behavior of the inflation rate differential, and the behavior of the nominal exchange rate changes.

Beginning with Mussa (1986), a large body of literature demonstrates that real exchange rate behavior crucially depends on the underlying nominal exchange rate regime. In fact, Mussa (1986) shows that real exchange rates are more volatile under floating compared to fixed nominal exchange rate regimes. Moreover, he finds that this extra variability comes primarily from the nominal exchange rate rather than the inflation differential. Ever since, a number of empirical and theoretical studies have found support for this argument (Baxter and Stockman, 1989; Dixon, 1999; Flood and Rose, 1999; Lothian and McCarthy, 2002; MacDonald, 1999). This is not surprising because under a floating regime, the nominal exchange rate represents an asset that is traded daily on financial markets. On the other hand, the inflation rate differential is one of the most important overall macroeconomic indicators, depending on numerous economy-wide factors, and does not show asset characteristics. Therefore, we are interested in the following question: to what extent does the value of the risk-sharing index depend on the nominal exchange rate regime?

Does the stochastic discount factor approach imply perfect risk-sharing between countries when their nominal exchange rate is fixed?

The main issue we want to investigate is whether by fixing the (bilateral) nominal exchange rate, two countries "*automatically (or mechanically)*" achieve perfect levels of international risk-sharing according to the approach developed in this study. In order to answer this question(s), we calculate risk-sharing indices for two episodes with fixed nominal exchange rate regimes. The first one refers to the experience of the countries in the EMU after the introduction of the common currency - the Euro - on 1 January 1999. The second example refers to a group of emerging economies that had their nominal exchange rates (almost) fixed against the US dollar over extended periods of time in recent past.

5.2.1 Eurozone

Data Sources

We focus on 12 countries from the Eurozone in the period January 1993-December 2005. We collect monthly data on widely-used stock market indices for each country and for the Eurozone as a whole. For the Eurozone we include two indices: Eurozone FTSE Local made up of major Eurozone-based corporations that receive more than 70 percent of their total revenues domestically, i.e. from within the Eurozone; and Dow Jones Euro Stoxx 50 - the most widely-used index of 50 major corporations based in the Eurozone. A list with stock market indices per country can be found in Table 5.1.[1]

Moreover, we collect data series on interest rates for one-month Eurocurrency deposits, consumer price indices, and nominal exchange rates for the pre-Euro period. Eurocurrency deposit rates for Greece and for the Eurozone were not available for the entire time period. Therefore, we supplemented the series for Greece for the period January 1993-January 2001 with interest rates series on bank deposits in domestic currency (Greek drachmas). For the Eu-

[1] The leading stock market indices for Greece, Italy and Luxembourg miss several observations for the period January 1993-December 2005. Therefore, they have been supplemented by alternative stock market indices: Athex Composite for ASE 20 (Greece), MIB Storico General for MIB 30 (Italy), and LUXX Datastream-Calculated for LUXX (Luxembourg).

Table 5.1: Stock Market Indices

Country	Index
AUS	Austrian Traded ATX
BEL	BEL 20
FIN	OMX Helsinki 25
FRA	CAC 40
GER	DAX 30
GRE	ASE 20
IRE	ISEQ Irish Overall
ITA	MIB 30
LUX	LUXX
NL	AEX
POR	Portugal PSI General
SPA	IBEX 35

Note: This table contains a list of stock market indices for the Eurozone countries. When available, we include the leading stock market indices for each of the 12 countries. Some of the missing observations in the leading indices for Greece, Italy, and Luxembourg have been supplemented by alternative stock market indices: Athex Composite for ASE 20 (Greece), MIB Storico General for MIB 30 (Italy), and LUXX Datastream-Calculated for LUXX (Luxembourg).

rozone we use the German interest rates on one-month Eurocurrency deposits in the period January 1993-December 1998. After the introduction of the Euro, we use the Eurozone rate on one-month Eurocurrency deposits.[2] CPI indices refer to mid-month values of the consumer price indices. Therefore, we calculate inflation rates as monthly changes in the CPI. For each country in the pre-Euro period we use nominal spot exchange rates (in terms of domestic currency units per European Currency Unit (ECU)) at the beginning of the month. Moreover, for Greece we use the same nominal spot exchange rate until June 2000, when the drachma was fixed against the Euro.

In order to capture the importance of the introduction of the Euro, i.e. the fixation of all bilateral nominal exchange rates, we separate the entire time period into two subperiods of approximately equal length: one before the

[2]Therefore, due to the convergence of interest rates after the introduction of the Euro in January 1999, one may argue that the Eurozone interest rate equals the German rate throughout the entire time period.

introduction of the Euro, from January 1993 till December 1998; and the other after the introduction of the Euro, from January 1999 till December 2005. In this way, we separate the dataset symmetrically with respect to 1999 - the year of the introduction of the Euro.[3] Subsequently, we do the calculations for each subperiod separately.

Summary Statistics

The summary statistics for the countries from the Eurozone are presented in Tables 5.2, 5.3, and 5.4. Table 5.2 shows annualized means and standard deviations for the real excess stock market returns for two sub-periods: before the introduction of the Euro (1993-1998), and after (1999-2005). The data suggests large differences between mean excess returns in the two subperiods: stock market returns in the pre-Euro period were much higher compared to the Euro period. To a large extent, this behavior is influenced by the "dot-com bubble", which was characterized by a prolonged upward trend in the first, and abrupt downward correction in the second subperiod. However, while the mean excess returns were clearly influenced by this bubble, their volatility was not: standard deviations are comparable across the two subperiods.

In fact, this joint development - a decrease in mean excess returns without a significant change in their variability, implies lower Sharpe ratios. Indeed, the corresponding columns in Table 5.2 indicate that the Sharpe ratio decreased for all but one Eurozone country (Austria). As the Sharpe ratio measures investor's compensation for risk taken, its declining trend suggests that markets lowered the reward given to risk-bearers in the second (Euro-period) compared to the first (pre-Euro) period.

Table 5.3 gives similar type of information for real foreign exchange excess returns. It shows the annualized means and standard deviations for real excess returns earned on the foreign exchange market.[4] Two important points deserve attention in this figure. First, mean real excess returns are typically very small in absolute value and of comparable magnitudes across time subperiods.

[3]Therefore, we control for the possibility that the number of observations might have an influence on the results from the calculations.

[4]ECU is the benchmark currency in the pre-Euro period, and the Euro is the benchmark currency in the calculations for the Euro-period.

Table 5.2: Summary Statistics: Stock Market Returns (Annualized)

	1993-1998			1999-2005		
country	mean	std dev	Sharpe ratio	mean	std dev	Sharpe ratio
AUS	8.72	23.04	0.38	18.25	16.37	1.11
BEL	13.93	16.19	0.86	-0.57	17.03	-0.03
FIN	29.65	24.76	1.19	10.94	28.82	0.38
FRA	8.91	21.02	0.42	2.25	20.06	0.11
GER	16.29	19.42	0.84	1.42	24.74	0.06
GRE	5.86	38.60	0.15	2.08	28.24	0.07
IRE	24.05	17.85	1.35	7.56	17.79	0.43
ITA	13.89	26.23	0.53	0.29	20.15	0.01
LUX	19.29	14.89	1.29	9.91	25.12	0.39
NL	19.51	18.37	1.06	-2.02	22.63	-0.09
POR	19.06	24.45	0.78	-1.28	16.39	-0.08
SPA	18.36	23.83	0.77	0.95	20.48	0.05
EURO	15.62	17.99	0.87	1.09	20.89	0.05

Note: This table contains summary statistics for stock market excess returns for each of the 12 countries from the Eurozone. The first three columns contain statistics for the pre-Euro period (January 1993-December 1998), while the last three columns contain statistics for the Euro-period (January 1999-December 2005). Excess returns are calculated over one-month Eurocurrency deposit rates for the corresponding country. The excess returns for the Eurozone are calculated over one-month Eurocurrency deposit rates for Germany (in the pre-Euro period) and for the Eurozone (in the Euro-period). All data-series are retrieved from Datastream and all summary statistics are annualized and expressed in percentage terms (rounded to two decimal places).

Second, the real foreign exchange excess returns show much lower volatility in the second (after the introduction of the Euro) compared to the first time subperiod (before the Euro). The latter observation is in line with the common intuition that (excess returns on) real exchange rates become much less volatile under fixed nominal exchange rate regimes.

Finally, Table 5.4 presents correlations between returns on stock market indices in the countries from the Eurozone and the overall Eurozone index Euro Stoxx 50 for the two time subperiods. We also used an alternative overall index for the Eurozone - Eurozone FTSE Local - but the correlations were very similar and we do not present them here.[5] The general impression

[5]Calculations based on this alternative index for the Eurozone differ only marginally.

Table 5.3: Summary Statistics: Foreign Exchange Returns (Annualized)

	1993-1998		**1999-2005**	
country	mean	std dev	mean	std dev
AUS	4.81	3.03	2.94	0.38
BEL	-0.35	2.42	0.04	0.17
FIN	4.96	6.15	2.59	0.57
FRA	-0.97	2.47	-0.31	0.13
GER	0.20	2.89	-0.56	0.15
GRE	-7.54	5.22	-1.63	1.35
IRE	3.73	5.56	4.60	0.63
ITA	-3.81	6.89	0.35	0.16
LUX	-0.31	2.45	0.24	0.17
NL	0.40	2.77	0.56	0.34
POR	-4.88	3.16	1.04	0.30
SPA	-4.49	5.01	1.12	0.16

Note: This table contains summary statistics for excess returns on the foreign exchange market for each of the 12 countries from the Eurozone. The first two columns contain statistics for the pre-Euro period (January 1993-December 1998), while the last two columns contain statistics for the Euro-period (January 1999-December 2005). Excess returns on the foreign exchange market are calculated as (real) deviations from uncovered interest rate parity ($\theta^e + r^f - r^d$). Eurozone is used as the benchmark (domestic) country in all calculations. ECU is the benchmark currency in the pre-Euro period, and the Euro is the benchmark currency in the Euro-period. All data-series are retrieved from Datastream and all summary statistics are annualized and expressed in percentage terms (rounded to two decimal places).

is that the correlations for most national indices with the Eurozone index were very high. Moreover, these stock market returns correlations increased after the introduction of the common currency for most countries. Austria, Belgium, and Portugal are exceptions to this trend because the stock market correlations for these countries (slightly) decreased with the introduction of the Euro.

Results for the Eurozone

The results for the risk-sharing index are presented in Table 5.5. All figures are calculated following the bilateral framework between each of the 12 Eu-

Table 5.4: Stock Market Correlations with Euro Stoxx 50

Country	**pre-Euro**	**Euro**
AUS	0.793	0.476
BEL	0.845	0.684
FIN	0.671	0.825
FRA	0.929	0.980
GER	0.914	0.944
GRE	0.358	0.531
IRE	0.637	0.699
ITA	0.720	0.867
LUX	0.558	0.717
NL	0.903	0.934
POR	0.781	0.638
SPA	0.823	0.833

Note: This table presents figures for stock market correlations for all 12 countries from the Eurozone. All correlations are measured by the coefficient of correlation between monthly returns on each stock market index and monthly returns on Euro Stoxx 50 - the overall index for the Eurozone. The first column contains statistics for the pre-Euro period (January 1993-December 1998), while the second column contains statistics for the Euro-period (January 1999-December 2005).All figures are rounded up to three decimal places.

rozone countries and the Eurozone as a whole.[6] The first column refers to the results before, and the second column to the results after the introduction of the common currency. Two important observations deserve attention in this table. First, the risk-sharing index gets very high values for the first subperiod, typically about 0.999. Second, the risk-sharing index in the second subperiod - after the introduction of the Euro - gets extreme values, even much higher than in the first subperiod. In fact, the risk-sharing index asymptotically approaches the value of 1 for each country in the Eurozone. Using the terminology in Brandt et al. (2006), these values suggest *perfect risk-sharing*

[6]Strictly speaking, data on the Eurozone should exclude the country concerned. In our case, since interest rates across the countries in the Eurozone are virtually the same and nominal exchange rates are fixed, this point can only be valid for the Eurozone inflation rate and Eurozone stock market returns. However, due to computational difficulties, we do not exclude the contribution of each country's inflation towards the Eurozone inflation rate nor the fact that some of the companies included in the Euro Stoxx 50 index are based in the corresponding countries.

Table 5.5: Risk-Sharing Index

Country	pre-Euro	Euro
AUS	0.999893	0.999999
BEL	0.999656	0.999975
FIN	0.999011	0.999999
FRA	0.999842	0.999999
GER	0.999502	0.999999
GRE	0.999635	0.999941
IRE	0.999448	0.999999
ITA	0.998154	0.999999
LUX	0.999842	0.999999
NL	0.999741	0.999998
POR	0.999841	0.999999
SPA	0.999276	0.999999

Note: This table presents figures for the bilateral risk-sharing index between each of the 12 Eurozone countries and the Eurozone as a whole. The risk-sharing index is calculated according to the following formula: $RSI = 1 - \frac{\Sigma^{ee}}{\mu^{d\prime}\Sigma^{-1}\mu^{d} + \mu^{f\prime}\Sigma^{-1}\mu^{f}}$. The first column contains statistics for the pre-Euro period (January 1993-December 1998), while the second column contains statistics for the Euro-period (January 1999-December 2005). All figures are rounded up to six decimal places.

between each of these 12 countries and the Eurozone, and thereby, between each individual pair of countries.

Understanding the Results

Having observed these extreme values for the international risk-sharing index, we now attempt to understand the underlying reasons. Therefore, we look at the behavior of the two index components: real exchange rate volatility, and discount factor volatility.

Figure 5.1 shows monthly changes in the nominal exchange rate (domestic currency per unit of ECU or Euro) for 12 countries from the Eurozone. The vertical axis measures monthly percentage changes in the exchange rate, and the vertical line in the middle of the figure denotes the separation of the complete time period into pre-Euro and Euro subperiods. The figure shows a continuous decrease in the nominal exchange rate variability for each country in the pre-Euro period. Moreover, it shows zero-variability in 11 out 12 nominal exchange rates starting in January 1999 with the adoption of the Euro.[7]

In order to see the link between nominal and real exchange rate volatility, we present similar graph for the real exchange rates in Figure5.2. This figure shows that the behavior of the real exchange rate was very similar to the behavior of the nominal exchange rate. In fact, the time-lines in Figure 5.2 closely resemble those for the nominal exchange rate in Figure 5.1. Therefore, the adoption of the Euro does not only eliminate nominal, but also the major (bulk) portion of real exchange rate volatility. There are only minor fluctuations left after January 1999, coming exclusively from the inflation rate differential. The only exception is the real exchange rate for Greece, which stabilizes with a two-year lag, due almost entirely to its late adoption of the Euro.

[7]The only exception is Greece because of its late adoption of the Euro. The Greek drachma was fixed against the Euro on 19 June 2000, and officially replaced by the common European currency on 1 January 2001.

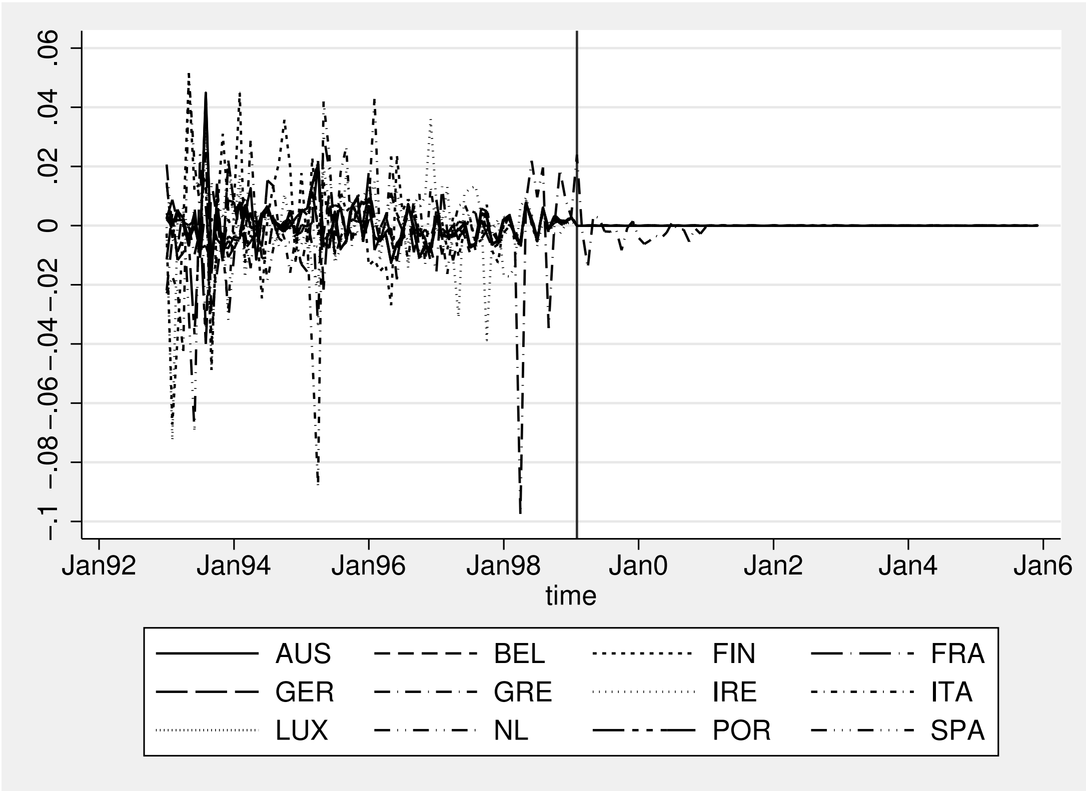


Figure 5.1: Nominal Exchange Rates (1993-2005)

Note: This figure presents monthly changes in the nominal exchange rates for all 12 countries in the Eurozone over the period January 1993-December 2005. The ECU is used as benchmark currency in the pre-Euro period, and the Euro afterwards. Positive changes in the nominal exchange rate indicate depreciation of the benchmark currency (appreciation of the currency for the corresponding country). The vertical line marks the introduction of the Euro (January 1999).

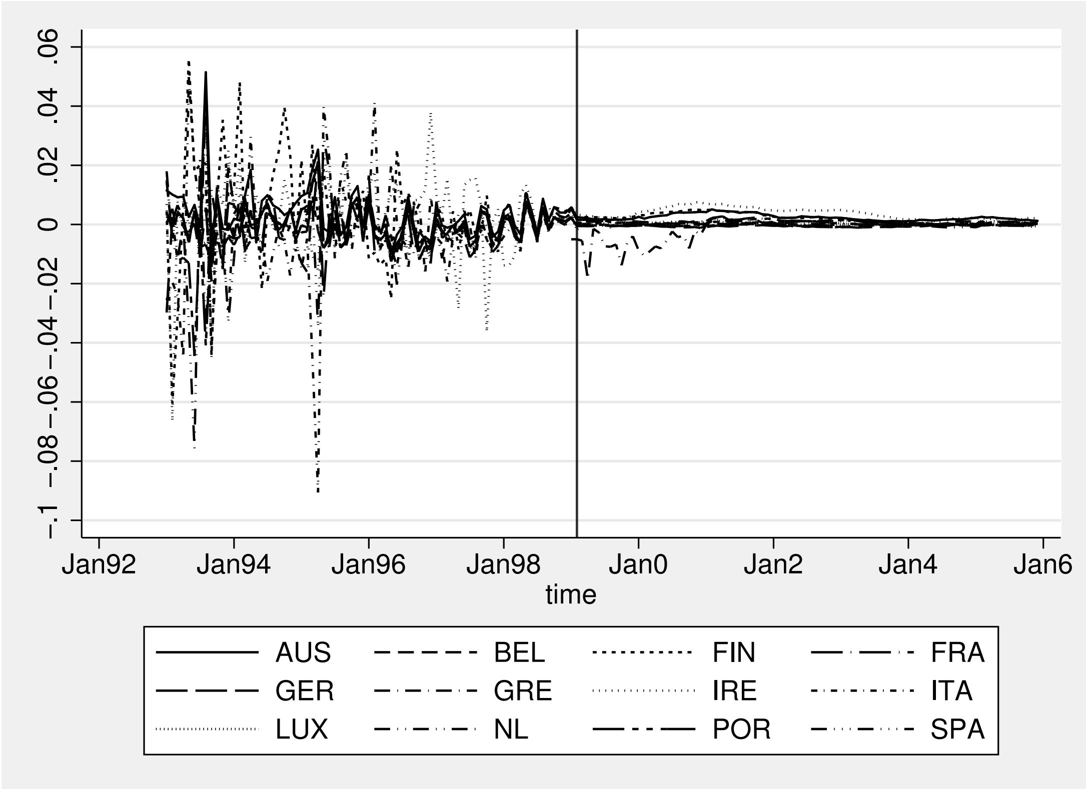


Figure 5.2: Real Exchange Rates (1993-2005)

Note: This figure presents monthly changes in the real exchange rates for all 12 countries in the Eurozone over the period January 1993-December 2005. The ECU is used as benchmark currency in the pre-Euro period, and the Euro afterwards. Real exchange rate changes are calculated as nominal exchange rate changes corrected for the inflation differential. Positive changes in the real exchange rate indicate real depreciation of the benchmark currency (real appreciation of the currency for the corresponding country). The vertical line marks the introduction of the Euro (January 1999).

Table 5.6: Real Exchange-Rate Volatility (Annualized Values)

Country	pre-Euro	Euro
AUS	3.01	0.38
BEL	2.40	0.17
FIN	6.11	0.57
FRA	2.45	0.13
GER	2.87	0.15
GRE	5.19	1.34
IRE	5.52	0.63
ITA	6.84	0.16
LUX	2.43	0.17
NL	2.75	0.34
POR	3.14	0.30
SPA	4.97	0.16

Note: This table presents figures for real exchange rate volatility for each of the 12 Eurozone countries. Volatility is measured by the annualized standard deviation calculated from monthly changes in the real exchange rate. The first column contains statistics for the pre-Euro period (January 1993-December 1998), while the second column contains statistics for the Euro-period (January 1999-December 2005). All figures are expressed in percentages terms, rounded up to two decimal places.

In Tables 5.6, 5.7, and 5.8 we go one step further in understanding the results for the risk-sharing index in the Eurozone. Tables 5.6 and 5.7 show statistics for real exchange rate and discount factor volatility in each country from the Eurozone. They refer to the annualized values of the two components used in the calculation of the risk-sharing index. Table 5.6 gives additional evidence about the rapid drop in the real exchange rate volatility after the adoption of the Euro. In fact, all figures in the second column (Euro period) are on average about 10 times lower than the corresponding entries in the first column (pre-Euro period).

Table 5.7 shows annualized values for the discount factor volatility in each country. The first columns in Table 5.7 give values for the Eurozone discount factor volatility and the last two columns give the corresponding values for the other countries used in the bilateral calculations. Moreover, for each country, the first column refers to the pre-Euro, while the second for the Euro period.

Table 5.7: Discount Factors Volatility (Annualized Values)

	Eurozone DF	Eurozone DF	Country DF	Country DF
Country	pre-Euro	Euro	pre-Euro	Euro
AUS	206.69	811.29	204.38	810.93
BEL	91.33	24.48	91.74	24.34
FIN	139.02	479.16	135.54	478.63
FRA	137.41	250.85	138.13	250.98
GER	90.94	375.39	90.76	375.54
GRE	189.96	123.65	193.96	124.96
IRE	167.38	754.88	165.23	754.27
ITA	110.80	223.01	114.39	222.86
LUX	136.44	150.46	136.69	150.31
NL	120.92	173.56	120.622	173.24
POR	173.93	359.45	176.74	359.16
SPA	128.83	713.94	132.37	713.79

Note: This table presents figures for discount factor volatility for each of the 12 Eurozone countries. Volatility is calculated by the annualized standard deviation of the discount factor according to equation (4.21): $\frac{1}{dt}\sigma^2(d\ln\Lambda^i) = \mu^{i\prime}\Sigma^{-1}\mu^i$. The first column contains statistics for the pre-Euro period (January 1993-December 1998), while the second column contains statistics for the Euro-period (January 1999-December 2005). All figures are expressed in percentage terms, rounded up to two decimal places.

Hence, we can see that the discount factors display similar volatility levels for each Eurozone-country pair (compare column 1 with column 3, or column 2 with column 4). Furthermore, for most countries, discount factors became more variable in the later period, although these changes are not as strong as in the case of the real exchange rate. In sum, the changes in both components of the risk-sharing index work in the same direction: they unambiguously lead to inflation of the index. In order to see the relative importance of each component, we proceed with a comparison of the changes in their volatility levels between the two subperiods.

Table 5.8 presents ratios between the volatility level in the pre-Euro period over the volatility level in the Euro period for each index component. Values above 1 indicate a decrease in the variability of the corresponding variable after the introduction of the Euro. We present ratios for the following five variables:

real exchange rate, Eurozone discount factor, country discount factor, sum of both discount factors, and for the ratio of the real exchange rate over the sum of the discount factors. The first column in Table 5.8 displays ratios much higher than 1, indicating huge drop in the real exchange rate volatility for each country in the Eurozone. Moreover, these ratios differ widely across different countries. Thus, the real exchange rate was about 8 times more volatile in the pre-Euro period for Austria (ratio of 7.87) and the Netherlands (8.05), but 31 times for Spain and 43 times for Italy. The smallest drop comes for Greece (ratio of 3.86), but this is largely due to the fact that in the period January 1999-June 2000 its nominal exchange rate was not fixed yet. The next three columns generally show values below 1, indicating an increase in the volatility of the discount factors (for the Eurozone, for the corresponding country, and for the sum of the two).[8] Finally, the last column presents changes in the overall ratio of real exchange rate volatility over the sum of discount factor volatilities. Therefore, this ratio shows the mirror image of the risk-sharing index (see equation (4.13)). Values higher than 1 indicate that this ratio was several times higher before the introduction of the Euro for each country in the dataset. Moreover, the decrease of this ratio (and the corresponding increase in the risk-sharing index) was not identical for all countries that adopted the Euro: it was just a little bit higher for Greece (2.50) and Belgium (3.73), while much higher for Italy (86.39) or Spain (172.78)in the pre-Euro period.[9] In sum, the relative volatility levels just strengthen the evidence from Tables 5.6 and 5.7: a strong decrease in real exchange rate volatility and a slight increase in (the sum of) discount factors volatility yielding an immense decrease in their ratio.

Unambiguously, the adoption of the Euro influenced both components of the risk-sharing index: directly, it led to a fall in real exchange rate volatility across Eurozone countries, and indirectly, it contributed to an increase in discount factor(s) volatility. While we elaborated broadly on the first effect

[8]The exceptions being Belgium and Greece, where stock market returns imply less volatile discount factors after the introduction of the Euro.

[9]The main reason for the very small change in the case of Greece is its late adoption of the Euro. For Belgium this is primarily due to its much calmer stock market index BEL 20 in the Euro period.

Table 5.8: Relative Volatility Levels (pre-Euro period/Euro period)

Country	X-Rate	Eurozone DF	Country DF	Sum of DFs	X-Rate/DFs
AUS	7.88	0.25	0.25	0.25	31.09
BEL	13.98	3.73	3.77	3.75	3.73
FIN	10.81	0.29	0.28	0.29	37.71
FRA	19.56	0.55	0.55	0.55	35.63
GER	18.78	0.24	0.24	0.24	77.62
GRE	3.86	1.54	1.55	1.54	2.50
IRE	8.78	0.22	0.22	0.22	39.82
ITA	43.63	0.49	0.51	0.51	86.39
LUX	14.61	0.90	0.91	0.91	16.08
NL	8.05	0.69	0.69	0.69	11.56
POR	10.63	0.48	0.49	0.49	21.79
SPA	31.61	0.18	0.19	0.18	172.78

Note: This table compares the volatility of the risk-sharing index components before and after the introduction of the Euro. It presents ratios for the volatility level in the pre-Euro period over the volatility level in the Euro period for each country in the Eurozone. Values above 1 indicate a decrease in the variability of the corresponding variable after the introduction of the Euro. The first three columns contain figures for real exchange rate volatility (Σ^{ee}), and discount factor volatilities ($\mu^{d\prime}\Sigma^{-1}\mu^{d}$ and $\mu^{f\prime}\Sigma^{-1}\mu^{f}$). The last two columns contain figures for the sum of discount factor volatilities and for the ratio between unshared risk and total risk in the two countries. All figures are calculated as ratios of the annualized standard deviations for the corresponding variables.

above, here we pay more attention to the latter effect. Therefore, we present the discount factor loadings (calculated as $\mu^d\Sigma^{-1}$ and $\mu^f\Sigma^{-1}$) in Table 5.9. This table contains three rows for each country: one for each source of shocks in the model. The first row refers to the discount factor loading on the Eurozone stock market shocks (domestic stock market shock), the second to the real exchange rate shock (between the Eurozone and the corresponding country), and the third to the stock market shock of the corresponding (foreign) country. In line with equation (4.25), domestic and foreign discount factors load equally on each of the stock market shocks, and domestic discount factor loads on the exchange rate shocks by one more than the foreign discount factor. The last point implies that the difference between the two discount factors at each point in time equals the exchange rate. Moreover, two findings in Table 5.9 deserve particular attention.

First, most loadings (especially those on exchange rate shocks) are higher (in absolute value terms) compared to those given in Brandt et al. (2006). Second, the loadings on the exchange rate shocks attain *extremely* high values in the second period (1999-2005). This latter finding means that both discount factors load much more on the exchange rate shocks relative to the stock market shocks. In fact, both of these findings come as a result of the extremely low volatility of the real exchange rate. Since the discount factors are calculated by the following formula $\mu^d\Sigma^{-1}$, low values for the denominator (variability of the real exchange rate) imply high values for whole term (discount factor loading). Thus, limited volatility of the real exchange rate explain why all discount factor loadings in the period preceding the introduction of the Euro are higher than in Brandt et al. (2006).[10] In turn, the extremely low real exchange rate variability is the main reason for the extremely high values for the exchange rate shock loadings after the introduction of the Euro. Finally, some of the discount factor loadings in Table 5.9 change sign. For example, some stock markets with relatively low Sharpe ratio are associated with negative discount factor loadings. This reflects the fact that certain markets are strictly dominated by (a combination of) other assets present in the system. In fact, investors that assets in some of these “dominated” markets forego better risk-return profiles (investment opportunities) available in other markets. Therefore, this sub-optimal (anomalous) behavior is reflected in the anomalous sign for the discount factor loadings on these assets (markets).[11]

5.2.2 Emerging Economies

In this section, we show the limitations of the stochastic discount factor approach to international risk-sharing under fixed nominal exchange rates with a second example. In the past two decades many developing or emerging countries adopted regimes of fixed or very rigid exchange rates against some of

[10] In fact, low variability for the real exchange rate implies low values for most terms in the variance-covariance matrix used in the calculation of all discount factors.

[11] This sub-optimal situation is not allowed in the equilibrium asset pricing model. In fact, the model suggests that shocks on all stock markets will be associated with positive loadings in equilibrium: favorable (negative) shock leads to a decrease (increase) in marginal utility growth. Clearly, this framework does not allow for dominated assets in equilibrium.

Table 5.9: Discount Factor Loadings: Eurozone Countries

		1993-1998		1999-2005	
Country		Eurozone	Country	Eurozone	Country
AUS	dz^d	11.23	11.23	1.73	1.73
	dz^e	60.92	59.92	2148.07	2147.07
	dz^f	-4.72	-4.72	14.76	14.76
BEL	dz^d	3.02	3.02	0.67	0.67
	dz^e	5.92	4.92	132.01	131.01
	dz^f	2.77	2.77	-1.10	-1.10
FIN	dz^d	2.03	2.03	-5.56	-5.56
	dz^e	11.51	10.51	856.79	855.79
	dz^f	3.49	3.49	7.77	7.77
FRA	dz^d	19.44	19.44	2.50	2.50
	dz^e	-4.45	-5.45	-2015.42	-2016.42
	dz^f	-13.56	-13.56	-1.14	-1.14
GER	dz^d	3.22	3.22	11.03	11.03
	dz^e	8.35	7.35	-2518.02	-2519.02
	dz^f	1.90	1.90	-7.26	-7.26
GRE	dz^d	7.48	7.48	-1.41	-1.41
	dz^e	-33.55	-34.55	-93.99	-94.99
	dz^f	-1.64	-1.64	0.80	0.80
IRE	dz^d	-1.44	-1.44	10.86	10.86
	dz^e	18.88	17.88	1240.52	1239.52
	dz^f	9.74	9.74	-2.69	-2.69
ITA	dz^d	4.26	4.26	0.02	0.02
	dz^e	-10.78	-11.78	1428.25	1427.25
	dz^f	1.04	1.04	-0.99	-0.99
LUX	dz^d	1.53	1.53	-1.96	-1.96
	dz^e	17.16	16.16	850.10	849.10
	dz^f	8.74	8.74	2.62	2.62
NL	dz^d	-4.41	-4.41	7.78	7.78
	dz^e	21.49	20.49	500.25	499.25
	dz^f	10.66	10.66	-6.49	-6.49
POR	dz^d	4.87	4.87	-0.72	-0.72
	dz^e	-48.29	-49.23	1243.58	1242.58
	dz^f	-0.50	-0.50	4.57	4.57
SPA	dz^d	4.57	4.57	-1.26	-1.26
	dz^e	-18.95	-19.95	4546.58	4545.58
	dz^f	0.51	0.51	-0.42	-0.42

Note: This table presents figures for the discount factor loadings in the bilateral setting. The loadings for the Eurozone (domestic) discount factor are given by $\mu^{d\prime}\Sigma^{-1}$ and the corresponding loadings for the Eurozone countries (foreign) discount factors are given by $\mu^{f\prime}\Sigma^{-1}$. The row marked dz^d contains figures for discount factor loadings on the Eurozone (domestic) stock market shocks, row dz^e refers to discount factor loadings on the foreign exchange market shocks, and row dz^f refers to discount factor loadings on the stock market shocks for the corresponding country in the Eurozone. The first two columns contain statistics for the pre-Euro period (January 1993-December 1998), while the last two columns contain statistics for the Euro-period (January 1999-December 2005).

the major world currencies, mainly against the dollar. The motivations for these moves differed from country to country, ranging from macroeconomic price stabilization, a desire to attract foreign capital, to promotion of more international trade. Common for all these cases is that both nominal and real exchange rates became much less variable against the major world currencies. Here we investigate whether these (quasi) fixed nominal exchange rate

Table 5.10: Stock Market Indices

Country	Index
ARG	BURCAP
CHI	IPSA (Indice de Precios Selectivo de Acciones)
KOR	KOSPI (Korea Composite)
MAL	KLCI (Kuala Lumpur Composite Index)
MEX	IPC 35 (Indice de Precios y Cotizaciones)
THA	SET 50 (Stock Exchange of Thailand)
UK	FTSE All
JAP	NIKKEI 225
USA	S&P 500

Note: This table contains a list of leading stock market indices used in the analysis for the emerging economies. The upper panel lists the stock market indices for the emerging economies, while the lower panel lists the stock market indices for UK, Japan, and USA.

episodes had a significant impact on the stochastic discount factor measures of international risk-sharing.

Data Sources

We collect monthly data on 6 emerging economies: Argentina, Chile, South Korea, Malaysia, Mexico, and Thailand in the period 1993-2005. Additionally, we collect data on Japan and UK for comparison. We use USA as the anchor country in this framework. Moreover, we separate the time period into two subperiods (1993-1998 and 1999-2005), exactly as in the case of the Eurozone. For the calculation of excess stock market returns, we take widely used stock market indices for the corresponding countries. A complete list of all stock market indices can be found in Table 5.10.[12]

Data on one-month Eurocurrency deposit rates for these emerging markets was not available in Datastream. Therefore, we include interest rates series on one-month domestic bank deposits to approximate for the risk-free rate.

[12] In case we cannot collect comparable data on the most widely used index for certain country, we include an alternative index. For example, we include BURCAP instead of MERVAL for Argentina.

Table 5.11: Summary Statistics: Stock Market Returns (Annualized)

	1993-1998			**1999-2005**		
country	mean	std dev	Sharpe ratio	mean	std dev	Sharpe ratio
ARG	9.06	34.08	0.27	18.44	44.80	0.41
CHI	0.49	26.49	0.02	9.26	17.37	0.53
KOR	-10.95	33.98	-0.32	12.81	35.22	0.36
MAL	-2.62	39.71	-0.07	6.94	26.78	0.26
MEX	-2.84	32.79	-0.09	19.20	27.58	0.69
THA	-15.76	44.11	-0.36	12.19	32.61	0.37
UK	5.49	12.67	0.43	-2.47	16.61	-0.15
JAP	-1.55	21.99	-0.07	2.09	21.70	0.10
USA	12.60	12.17	1.03	-1.19	17.01	-0.07

Note: This table contains summary statistics for stock market excess returns for emerging economies and industrial countries. The upper panel refers to the group of emerging economies (Argentina, Chile, Korea, Malaysia, Mexico, and Thailand), while the lower panel refers to the group of industrial countries (UK, Japan, and USA). The first three columns contain statistics for period January 1993-December 1998, while the last three columns contain statistics for period January 1999-December 2005. Excess returns are calculated over one-month domestic bank deposit rates for the emerging economies, and over one-month Eurocurrency deposit rates for the industrail countries. All data-series are retrieved from Datastream and all summary statistics are annualized and expressed in percentage terms (rounded to two decimal places).

Nominal exchange rate are measured as domestic currency units per US dollar. Finally, inflation data comes from the CPIs of the corresponding countries measured at mid-month. For Japan, UK and USA we use the same data sources as in Brandt et al. (2006).

Summary Statistics

The summary statistics are presented in Tables 5.11 and 5.12. The figures in Table 5.11 demonstrate that stock market returns in emerging economies are much more volatile than in industrial countries. In fact, the annualized standard deviations for emerging economies (columns 2 and 4) are typically double the figures for industrial economies.[13] Moreover, columns 1 and 3 make clear that several stock market downfalls happened in this period. The effect of the financial crises in the second half of the 1990s is evident from this table.

[13] The only exception being Chile, where the figures are comparable to those in the industrial countries.

Table 5.12: Summary Statistics: Foreign Exchange Returns (Annualized)

	1993-1998		**1999-2005**	
country	mean	std dev	mean	std dev
ARG	-3.30	0.84	-17.43	27.15
CHI	-8.28	5.61	-2.37	9.79
KOR	-6.99	19.25	0.95	7.98
MAL	-5.90	15.57	-0.65	0.52
MEX	-15.68	19.36	1.46	8.85
THA	-6.45	18.47	-1.19	7.45
UK	0.31	7.35	-0.71	6.87
JAP	3.05	14.04	0.90	10.67

Note: This table contains summary statistics for excess returns on the foreign exchange market for the group of emerging economies and industrial countries. USA is used as the benchmark (domestic) country, and the US dollar is the benchmark currency in all calculations. Excess returns on the foreign exchange market are calculated as (real) deviations from uncovered interest rate parity ($\theta^e + r^f - r^d$): borrowing at the US interest rate, converting to the foreign currency, investing on the foreign interest rate, and converting the proceeds back to US dollars. The first two columns contain statistics for the period January 1993-December 1998, while the last two columns contain statistics for the period January 1999-December 2005. All data-series are retrieved from Datastream and all summary statistics are annualized and expressed in percentage terms (rounded to two decimal places).

In fact, most emerging economies display negative values for real excess stock market returns (columns 1 and 3) in the first subperiod. A similar conclusion applies to the industrial countries in the second period, reflecting the burst of the dotcom bubble in 2000-2001. Finally, the Sharpe ratios closely follow the trend in mean excess returns, i.e. they increase for all emerging markets and Japan in the second period and decrease for the UK and the USA.[14]

Table 5.12 displays summary statistics for excess returns on the foreign exchange market. The negative values for most emerging economies in columns 1 and 3 indicate a net loss on investments in foreign exchange in these countries.[15] Moreover, excess returns volatility levels differ widely across countries and time subperiods. In general, foreign exchange excess returns vary much more in emerging economies compared to Japan or the UK. However, there

[14]Hence, the changes in the Sharpe ratios largely reflect the two stock market downturns in the corresponding periods: the emerging market crises in the first period and the dotcom-bubble in the second period.

[15]In fact, this partly reflects some peso-problem related issues.

Table 5.13: Stock Market Correlations with S&P 500

Country	1993-1998	1999-2005
ARG	0.549	0.272
CHI	0.480	0.488
KOR	0.238	0.677
MAL	0.371	0.299
MEX	0.481	0.666
THA	0.371	0.469
UK	0.682	0.823
JAP	0.404	0.501

Note: This table presents stock market correlations for the group of emerging and industrial countries. All correlations are measured by the coefficient of correlation between monthly returns on each stock market index and monthly returns on S&P 500. The first column contains statistics for the period January 1993-December 1998, while the second column contains statistics for the period January 1999-December 2005.All figures are rounded up to three decimal places.

are subperiods for certain emerging economies when exactly the opposite is true. For example, Argentina and Chile in the first subperiod, and Malaysia in the second subperiod display extremely low volatility in foreign exchange returns.

Finally, Table 5.13 presents stock market return correlations with the USA (S&P 500). The correlation coefficients for emerging economies are comparable to those for industrial countries. In fact, most stock markets correlations in emerging economies fall inbetween the values for Japan and UK.

Results for Emerging Economies

The results for the (bilateral) risk-sharing index between the corresponding countries and the USA are displayed in Table 5.14. Several observations from this table deserve attention. First, all values are very high, suggesting almost perfect levels of international risk-sharing of all countries with the USA. More precisely, there are three episodes of literally perfect risk-sharing: Argentina and Chile in the first period and Malaysia in the second period. All three episodes reflect the fact that nominal exchange rates were fixed against the US dollar in the corresponding periods. Second, emerging economies achieve

Table 5.14: Risk-Sharing Index

Country	1993-1998	1999-2005
ARG	0.999997	0.968468
CHI	0.999588	0.992272
KOR	0.988865	0.993513
MAL	0.991606	0.999966
MEX	0.992549	0.997553
THA	0.990011	0.993846
UK	0.997864	0.973277
JAP	0.992833	0.884481

Note: This table presents figures for the bilateral risk-sharing index between each country and USA. The risk-sharing index is calculated according to the following formula: $RSI = 1 - \frac{\Sigma^{ee}}{\mu^{d\prime}\Sigma^{-1}\mu^{d}+\mu^{f\prime}\Sigma^{-1}\mu^{f}}$. USA is the benchmark (domestic) country in all calculations. The first column contains statistics for the period January 1993-December 1998, while the second column contains statistics for the period January 1999-December 2005. All figures are rounded up to six decimal places.

similar or even higher levels of risk-sharing with the USA than Japan or the UK. In fact, the risk-sharing index for Argentina and Chile is higher, while for the other emerging economies it is very similar to the index for Japan or the UK in the first subperiod. In the second subperiod, the risk-sharing index for almost every emerging economy is higher than the index for Japan or the UK.[16] Hence, the stochastic discount factor approach suggests that emerging markets achieved more international risk-sharing with the USA, than the much more globally integrated economies like Japan and the UK. In sum, the results from this analysis imply that the nominal exchange rate regime has a more profound impact on the risk-sharing measure than do all other economic characteristics of the corresponding country-pairs.

Understanding the Results

The results for the risk-sharing index in Table 5.14 convey a very counterintuitive message. Despite the widely documented observation that Japan and the UK are more financially and economically integrated with the USA than the

[16] The only exception is Argentina whose risk-sharing index is slightly lower than that for the UK.

group of emerging economies, the stochastic discount factor approach suggests the opposite. In fact, we consider this result to be a major limitation of the framework presented in Brandt et al. (2006) and in this chapter. In order to understand the reasons underlying this result, we analyze the behavior of the two risk-sharing index components in the rest of this section.

Figures 5.3 and 5.4 show the behavior of the nominal and real exchange rates for the 6 emerging economies, respectively. The vertical axis measures monthly percentage changes in the exchange rate, while the vertical line in the middle of the figure denotes a separation of the complete time period into two subperiods. In general, these economies enjoyed periods of relatively stable currencies, characterized by very small fluctuations in their nominal exchange rates against the US dollar, which were interrupted by several crises. Three main crisis episodes, accompanied by sharp exchange rate fluctuations, can be seen in this figure: the Mexican peso crisis in 1994-1995, the East Asian financial crisis in the second half of 1997 (affected Thailand, Korea, and Malaysia), and the Argentine financial crisis in late 2001 - early 2002. As can be seen in the figure, each of these crises was characterized by large falls in the value of the domestic currency (sometimes followed by smaller reversals) against the US dollar. Furthermore, Figure 5.4 shows the behavior of the real exchange rate, measured as monthly changes in nominal exchange rates corrected for the inflation differential. As already documented in the example with the Eurozone, Figure 5.4 shows that nominal and real exchange rates are very closely related. In fact, the evidence in these figures only strengthens the argument that the variability of the real exchange rate crucially depends on the behavior of the nominal exchange rate.

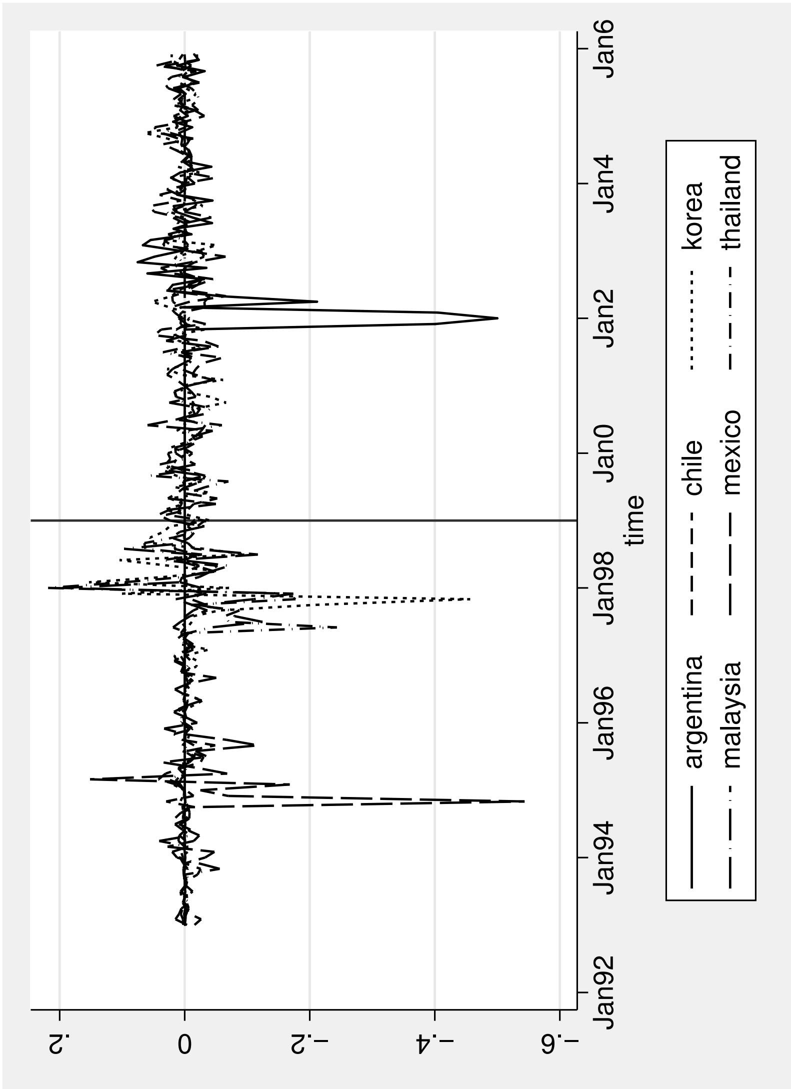


Figure 5.3: Nominal Exchange Rates (1993-2005)

Note: This figure presents monthly changes in the nominal exchange rates for the group of emerging and industrial countries over the period January 1993-December 2005. The US dollar is used as benchmark currency. Positive changes in the nominal exchange rate indicate depreciation of the US dollar (appreciation of the currency for the corresponding country). The separation into two time subperiods is marked by a vertical line (January 1999).

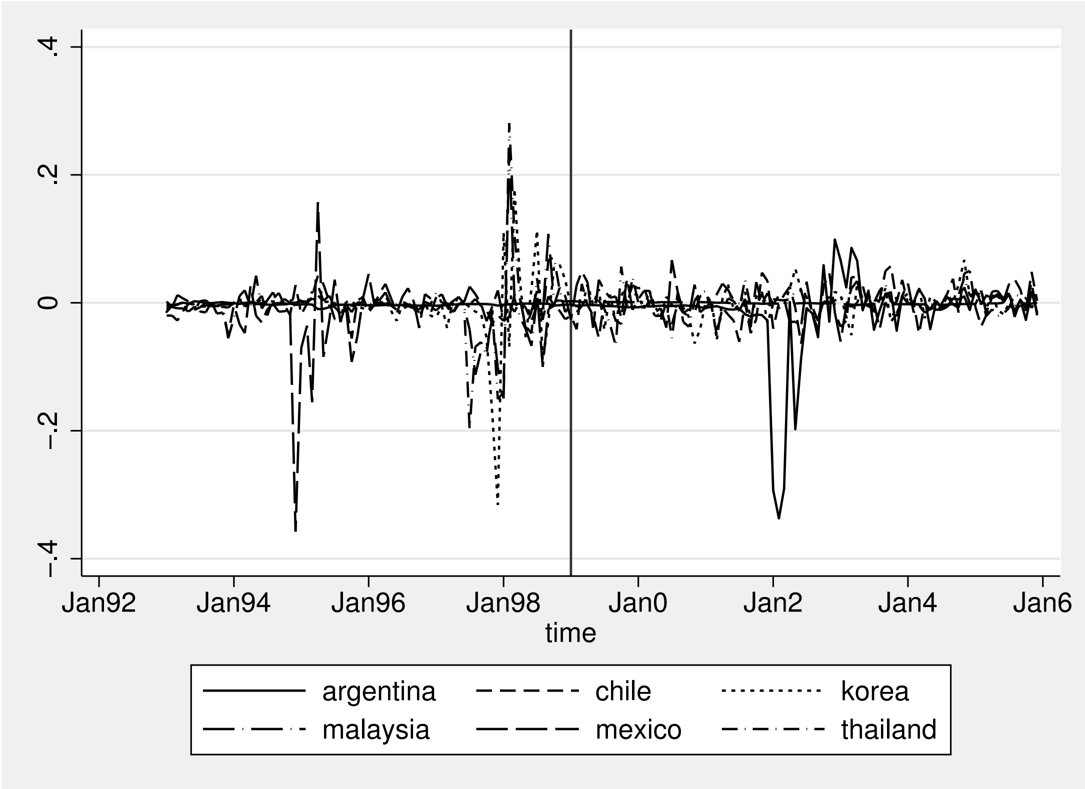


Figure 5.4: Real Exchange Rates (1993-2005)

Note: This figure presents monthly changes in the real exchange rates for the group of emerging and industrial countries over the period January 1993-December 2005. The US dollar is used as benchmark currency. Real exchange rate changes are calculated as nominal exchange rate changes corrected for the inflation differential. Positive changes in the real exchange rate indicate real depreciation of the US dollar (real appreciation of the currency for the corresponding country). The separation into two time subperiods is marked by a vertical line (January 1999).

Table 5.15: Real Exchange-Rate Volatility (Annualized Values)

Country	1993-1998	1999-2005
ARG	0.84	22.68
CHI	5.57	9.86
KOR	19.16	7.57
MAL	15.46	0.78
MEX	19.22	7.87
THA	18.37	6.60
UK	7.30	7.76
JAP	13.94	10.29

Note: This table presents figures for real exchange rate volatility for the group of emerging and industrial countries. Volatility is measured by the annualized standard deviation calculated from monthly changes in the real exchange rate against the US dollar (US dollar is used as benchmark currency). The first column contains statistics for the period January 1993-December 1998, while the second column contains statistics for the period January 1999-December 2005. All figures are expressed in percentages terms, rounded up to two decimal places.

Having visualized the relation between nominal and real exchange rate volatility, we turn to an in-depth analysis of the two components of the risk-sharing index here. Table 5.15 shows the real exchange rate volatility figures for each country in the dataset. All figures refer to annualized standard deviations measured in percentage terms over the corresponding subperiod.

In general, real exchange rate volatility decreased for most emerging economies in the second compared to the first time subperiod, reaching even lower levels than for Japan or the UK. The only exceptions are Argentina and Chile. In fact, the former experienced substantially higher real exchange volatility in the second subperiod mainly due to the currency crisis in 2001-2002 and the subsequent abandonment of the fixed one-to-one Argentine peso - US dollar parity. Moreover, two extremely low figures deserve special attention: Argentina in the first and Malaysia in the second subperiod. Both of these figures refer to fixed exchange rate episodes: the Argentine peso was pegged to the US dollar from 1991 till December 2001, and the Malaysian ringgit was fixed against the US dollar in the aftermath of the East Asian financial crisis in 1998. In turn, exactly these fixed exchange rate episodes correspond with the perfect risk-sharing scores in Table 5.14.

Table 5.16: Discount Factor Volatility (Annualized Values)

	USA DF	**USA DF**	**Country DF**	**Country DF**
Country	1993-1998	1999-2005	1993-1998	1999-2005
ARG	399.98	78.45	400.81	100.77
CHI	192.15	77.46	196.49	81.06
KOR	124.60	66.96	131.49	65.96
MAL	116.31	94.45	122.26	95.14
MEX	151.79	112.95	162.93	111.92
THA	126.63	58.32	132.89	60.69
UK	111.70	32.04	111.66	35.04
JAP	117.34	22.26	115.56	20.52

Note: This table presents figures for discount factor volatility for each of the group of emerging and industrial countries. Volatility is calculated by the annualized standard deviation of the discount factor according to equation (4.21): $\frac{1}{dt}\sigma^2(d\ln\Lambda^i) = \mu^{i\prime}\Sigma^{-1}\mu^i$. USA is used as benchmark (domestic) country in all calculations. The first column contains statistics for the period January 1993-December 1998, while the second column contains statistics for the period January 1999-December 2005. All figures are expressed in percentage terms, rounded up to two decimal places.

In Table 5.16 we present figures for the second component of the risk-sharing index - discount factor volatility. All results are based on bilateral calculations, where the first two columns refer to the anchor (domestic) country (USA for each country pair in this case), and the last two columns to the other country. As can be seen from the comparison of the results in columns 2 and 4 relative to columns 1 and 3, discount factor volatility decreased for most countries in the second compared to the first subperiod.

Finally, equation (4.13) indicates that the risk-sharing index changes if either the real exchange rate volatility or the discount factor volatility changes. In order to see the relative contribution of these two components for the risk-sharing index, we present comparisons of their values for the two subperiods in Table 5.17. This table presents ratios of volatility in the first subperiod over volatility in the second subperiod. Hence, values below 1 indicate volatility increase in the second compared to the first subperiod. As there was no systematic change comparable to the adoption of the Euro in the previous example, the values in this figure do not suggest any clear trend in real exchange rate volatility. In fact, it increased for two, decreased for the other

Table 5.17: Relative Volatility Levels (1993-1998 period/1999-2005 period)

Country	X-Rate	USA DF	Country DF	Sum of DFs	X-Rate/DFs
ARG	0.037	5.10	3.98	4.47	0.01
CHI	0.56	2.48	2.42	2.45	0.23
KOR	2.53	1.86	1.99	1.93	1.31
MAL	19.93	1.23	1.28	1.26	15.84
MEX	2.44	1.34	1.46	1.40	1.75
THA	2.78	2.17	2.19	2.18	1.27
UK	0.94	3.49	3.19	3.33	0.28
JAP	1.35	5.27	5.63	5.44	0.25

Note: This table compares the volatility of the risk-sharing index components between the first and the second subperiod. It presents ratios for the volatility level in the period 1993-1998 over the volatility level in the period 1999-2005. Values above 1 indicate a decrease in the variability of the corresponding variable in the latter period. The first three columns contain figures for real exchange rate volatility (Σ^{ee}), and discount factor volatilities ($\mu^{d\prime}\Sigma^{-1}\mu^{d}$ and $\mu^{f\prime}\Sigma^{-1}\mu^{f}$). The last two columns contain figures for the sum of discount factor volatilities and for the ratio between unshared risk and total risk in the two countries. USA is used as benchmark (domestic) country in all calculations. All figures are calculated as ratios of the annualized standard deviations for the corresponding variables.

four emerging markets, and remained stable for Japan and especially for the UK. However, two extreme figures deserve attention: first, there was a sharp increase in Argentine real exchange rate volatility, and second, Malaysian real exchange rate was 20 times higher in the first compared to the second period. Finally, the three middle columns show a clear trend in discount factor volatility: they became less variable for each country pair in the second subperiod. This reflects the fall in the highest Sharpe ratio shock documented in Table 5.11.[17]

We close the discussion of the results from the risk-sharing index with a short note on the discount factor loadings in Table 5.18. Similar as in the case with the Eurozone, this table contains three rows for each country: one for each source of shocks in the model. The first row refers to the discount factor loading on the domestic stock market shock (USA for all countries

[17]Discount factor volatility crucially depends on the excess return shock with the highest Sharpe ratio. Although the Sharpe ratios for all emerging markets increased in the second period, discount factor volatility decreased because of the dramatic fall in the highest Sharpe ratio from the first period - the US stock market shock (see Table 5.11).

here), the second to the real exchange rate shock (between the USA and the corresponding country), and the third to the (foreign) stock market shock of the corresponding country. In line with equation (4.25) and the evidence for the Eurozone, domestic and foreign discount factors load equally on each of the stock market shocks, and domestic discount factor loads on the exchange rate shocks by one more than the foreign discount factor.

Finally, two observations deserve particular attention. First, the discount factor loadings on the US stock market change sign in most cases:[18] from positive in the first they turn negative in the second period. This reflects the dramatic fall in the Sharpe ratio on the US stock market (the negative excess returns earned on the US stock market in the second period), which becomes "dominated" by the emerging markets' Sharpe ratios in second period. Second, the exchange rate loadings for Argentina and Chile in the first, and for Malaysia in the second subperiod obtain extremely high (absolute) values. In line with the discussion for the Eurozone, these loadings reflect rigidity in the corresponding exchange rates.

5.3 Discussion

In this study, we showed that the risk-sharing index leads to absurdly high results when applied to countries with fixed nominal exchange rate regimes. The main point to show here was that a large "disconnect" between the asset side (denominator) and the macroeconomic side (numerator) drives the results of the risk-sharing index. In fact, we showed that just by elimination of the asset component in the real exchange rate, the stochastic discount factor "automatically" leads to absurdly high levels of risk-sharing. First, we demonstrated that the stochastic discount factor model implies literally perfect risk-sharing across all countries from the Eurozone after the introduction of the Euro. This result reflected the nominal exchange rate regime and not other economic similarities or the high level of economic integration among this set of countries: the comparison with the pre-Euro period demonstrated

[18]The loading on the US stock market stays positive only in the calculations against the UK because the UK stock market has even more negative Sharpe ratio for the second period (see Table 5.11).

Table 5.18: Discount Factor Loadings: Emerging Economies

		1993-1998		1999-2005	
Country		USA	Country	USA	Country
ARG	dz^d	6.27	6.27	-0.73	-0.73
	dz^e	-466.85	-467.85	-3.13	-4.13
	dz^f	-2.01	-2.01	0.40	0.40
CHI	dz^d	11.40	11.40	-2.82	-2.82
	dz^e	-27.40	-28.40	-3.11	-4.11
	dz^f	-2.30	-2.30	5.15	5.15
KOR	dz^d	10.10	10.10	-4.84	-4.84
	dz^e	-2.08	-3.08	2.39	1.39
	dz^f	-1.45	-1.45	2.83	2.83
MAL	dz^d	9.68	9.68	-1.91	-1.91
	dz^e	-1.81	-2.81	-113.80	-114.80
	dz^f	-1.12	-1.12	1.89	1.89
MEX	dz^d	11.76	11.76	-7.05	-7.05
	dz^e	-5.19	-6.19	1.23	0.23
	dz^f	-1.10	-1.10	6.08	6.08
THA	dz^d	10.90	10.90	-2.45	-2.45
	dz^e	-0.12	-1.12	-3.68	-4.68
	dz^f	-1.74	-1.74	2.14	2.14
UK	dz^d	11.77	11.77	2.65	2.65
	dz^e	-0.52	-1.52	-3.13	-4.13
	dz^f	-4.48	-4.48	-4.03	-4.03
JAP	dz^d	10.48	10.48	-1.40	-1.40
	dz^e	0.50	-0.50	1.22	0.22
	dz^f	-2.65	-2.65	1.09	1.09

Note: This table presents figures for the discount factor loadings in the bilateral setting. The loadings for the USA (domestic) discount factor are given by $\mu^{d\prime}\Sigma^{-1}$ and the corresponding loadings for the other (foreign) discount factors are given by $\mu^{f\prime}\Sigma^{-1}$. The row marked dz^d contains figures for discount factor loadings on the USA (domestic) stock market shocks, row dz^e refers to discount factor loadings on the foreign exchange market shocks, and row dz^f refers to discount factor loadings on the stock market shocks for the corresponding (emerging or industrial) country. The first two columns contain statistics for the period January 1993-December 1998, while the last two columns contain statistics for the period January 1999-December 2005.

that the complete elimination of the nominal exchange rates had a profound effect for countries that already had very high risk-sharing indices. Second, we further strengthened this evidence with the analysis for the group of emerging markets. Irrespective of many underlying differences between the countries considered, fixed nominal exchange rates "automatically" implied perfect levels of international risk-sharing.

5.4 Concluding Remarks

In this study we present some new evidence of the stochastic discount factor approach to international risk-sharing. At the beginning, we present the theoretical framework that links the minimum-variance discount factors in two countries with the corresponding real exchange rate. We elaborate on the calculation of the discount factors and the construction of the risk-sharing index. Subsequently, we investigate the importance of the nominal exchange rate regime for the risk-sharing measures.

We arrive at two main conclusions: first, the international risk-sharing index calculated in this way crucially depends on the behavior of the nominal exchange rate. In fact, we demonstrate that this approach "mechanically" leads to (almost) perfect international risk-sharing under fixed nominal exchange rate regimes. This is true for all countries in the Eurozone, as well as the emerging economies like Argentina and Malaysia that had (almost) fixed exchange rates against the US dollar. Second, this approach suggests more integration in terms of better risk-sharing between several emerging economies and the USA than between the UK and the USA. Having demonstrated that this counterintuitive result is almost entirely driven by the rigidity in the nominal exchange rate, we consider it as a major limitation. In fact, the over-dependence of this approach on the nominal exchange rate regime restricts its application in cross-country risk-sharing comparisons.

In sum, the evidence in this study demonstrates that fixing the nominal exchange rate directly implies (almost) fixed real exchange rates, and, according to the stochastic discount factor model, perfect international risk-sharing as well. We conclude that real exchange rates might be very smooth (as in this case), but this does not necessarily imply perfect risk-sharing across countries.

Chapter 6

Does the Nominal Exchange Rate Explain the Backus-Smith Puzzle? Evidence from the Eurozone

6.1 Introduction

Efficient international risk-sharing implies equalization of the marginal utility growth rates across countries at any point in time. Moreover, using standard utility function of the constant relative risk aversion (CRRA) form, this condition implies equalization of consumption growth rates across countries. Then, relative differences in consumption growth between two countries should be unpredictable and totally independent of their bilateral real exchange rate. In fact, the real exchange rate does not play any role in this stylized world: since purchasing power parity is assumed to hold at any point in time, real exchange rates do not move at all.

Backus and Smith (1993) extend the basic real business cycle model by introducing non-traded goods, and therefore, the possibility that purchasing power parity does not hold. In turn, this extension modifies the general international risk-sharing condition: instead of having complete equalization of real marginal rates of intertemporal substitution (marginal utility growth rates), real exchange rate fluctuations act as a wedge between marginal utility growth rates in the latter model. Therefore, the marginal utility growth difference should be equal to the changes in the bilateral real exchange rate for the corresponding countries.[1] Furthermore, this condition suggests that countries with an appreciating currency (in real terms) should also experience relatively higher marginal utility growth. Finally, if consumer preferences in both countries can be represented by the power utility function of the constant relative risk-aversion (CRRA) class, then this condition has very clear, intuitive implications for consumption streams in both countries: consumption growth should be higher in countries that experience a relative drop in relative price of their goods, i.e. a real depreciation of their currency.

While the framework presented in Backus and Smith (1993) provides a plausible theoretical reason why consumption growth rates across countries are not very strongly correlated, it generates an additional empirical puzzle. In the same article, they show that data for OECD countries clearly reject their proposition: the relationship between real exchange rates and relative consumption growth is either insignificant, or more often, significantly negative. Actually, this implies an anomaly: consumption growth is relatively higher for those OECD countries that experience a relative increase in the overall price level. This finding is known in the literature as the *Backus-Smith puzzle* or the *consumption-real exchange rate correlation puzzle*.

The existence of the Backus-Smith puzzle in data for OECD countries has become an established finding in the international macroeconomics literature (Backus and Smith, 1993; Kollmann, 1995; Ravn, 2001). Not surprisingly then, most recent studies take it as given and try to develop theoretical models that produce outcomes matching the data, i.e. implying a negative correlation between real exchange rates and relative consumption growth. These models introduce various frictions into the canonical model(s) ranging from incomplete financial markets with non-traded goods sector (Benigno and Thoenissen, 2005), incomplete capital markets and distribution services that are intensive in local (domestic) inputs (Corsetti, Dedola and Leduc, 2004, 2007), economies with limited commitment (Kang, 2006), or private information (Kang, 2007).

[1] For derivation of this condition using an asset pricing framework, see Brandt et al. (2006), for example.

One of the rare recent empirical explanations for the Backus-Smith puzzle is presented by Hess and Shin (2006). Using data for all OECD countries, they separate real exchange rate movements into their underlying components: nominal exchange rate and inflation differential changes. They find that the nominal exchange rate is the main source of the Backus-Smith puzzle: while the correlation between the negative of the inflation differential and relative consumption growth is generally positive (as expected by theory), it is typically negative between the nominal exchange rate and relative consumption growth. Furthermore, this anomaly (negative consumption-real exchange rate correlation) is present primarily for episodes of very large nominal exchange rate fluctuations. Finally, using inter-state data for the USA, where the nominal exchange rate is constant, they show that the Backus-Smith puzzle disappears. In this way, the current chapter is very closely related to Hess and Shin (2006). Focusing on the Eurozone after the introduction of the Euro, it investigates the importance of (the elimination) of nominal exchange rate changes for the Backus-Smith anomaly.

Using data for the Eurozone countries, this study tries to give answers to the following questions: Is the nominal exchange rate the main reason for the consumption-real exchange rate (Backus-Smith) anomaly? Does the elimination of the nominal exchange rate solve the Backus-Smith puzzle?

The main findings in this study suggest that the nominal exchange rate is indeed the main reason for the anomalous correlation between relative consumption growth and real exchange rate changes. First, the puzzle disappears for the countries from the Eurozone after the introduction of the common currency. Second, the evidence from alternative floating (nominal) exchange rate periods suggests that only the nominal exchange rate displays anomalous behavior, while the inflation differential behaves in line with theory.

The rest of this chapter is organized as follows. Section 6.2 presents the theoretical framework and derives empirically testable versions of the main conditions. Section 6.3 describes the dataset and presents the empirical strategy. The results from the empirical analysis are presented in section 6.4, and finally, section 6.5 gives some concluding remarks.

6.2 Theoretical Framework

6.2.1 Backus-Smith Puzzle

Efficient resource allocation in an international context implies equalization of all marginal utility growth rates at each point in time.[2] Translated in a two-country framework, this condition suggests the following relation:[3]

$$\frac{u'(c_{t+1})}{u'(c_t)} = \frac{u'(c^*_{t+1})}{u'(c^*_t)}, \forall t \tag{6.1}$$

where $u'(.)$ stands for marginal utility ($u(.)$ is the corresponding utility function) and c_t stands for real consumption per capita at time t. In fact, this condition holds as long as all relative international prices stay constant, i.e. as long as the purchasing power parity (PPP) condition holds. As first pointed out by Backus and Smith (1993), deviations from PPP (due to the presence of non-tradeable goods, high transport costs or tariff barriers for example) drive a wedge between the marginal utility growth rates across countries.[4] In that case, Backus and Smith (1993) demonstrate that marginal utility growth rates have to be corrected for changes in relative prices, i.e. movements of the real exchange rate. Accordingly, the internationally efficient allocation will then be given by the following condition:[5]

$$\frac{e_{t+1}}{e_t}\frac{u'(c_{t+1})}{u'(c_t)} = \frac{u'(c^*_{t+1})}{u'(c^*_t)} \tag{6.2}$$

where e_t, the real exchange rate at time t, is defined as the price of foreign in terms of domestic goods:

$$e_t = s_t \frac{P^*_t}{P_t} \tag{6.3}$$

[2]In general, the utility function depends on many arguments and not only on the consumption level. The simplification made here emphasizes only the argument that takes central place in the empirical investigation.

[3]In the rest of the chapter, foreign variables are always denoted with a star (*).

[4]About evidence on the deviations from the purchasing power parity (PPP) condition see Engel (2000), Engel and Rogers (1996), Froot and Rogoff (1995), or Rogoff (1996), for example.

[5]For a formal, theoretical underpinning of this condition, see Backus and Smith (1993), Kollmann (1995), Ravn (2001), Hess and Shin (2006), or Corsetti et al. (2007), among others.

This equation clearly shows that the real exchange rate contains two components: the nominal exchange rate s_t which gives the price of foreign in terms of domestic currency (units of domestic currency per one unit of foreign currency) and the relative price component defined as the ratio of the CPIs for the two countries (foreign over domestic price level). Using this definition for the real exchange rate, condition (6.2) can be rewritten as follows:

$$\frac{s_{t+1}}{s_t}\frac{P^*_{t+1}}{P^*_t}\frac{P_t}{P_{t+1}}\frac{u'(c_{t+1})}{u'(c_t)} = \frac{u'(c^*_{t+1})}{u'(c^*_t)} \tag{6.4}$$

after a slight rearrangement this condition can be given in terms of rates of change:

$$\Delta u'(c^*_{t+1}) - \Delta u'(c_{t+1}) = \Delta s_{t+1} + \Delta p^*_{t+1} - \Delta p_{t+1} \tag{6.5}$$

where Δx_{t+1} denotes the rate of change of variable x between time $t+1$ and t, and Δp_{t+1} and Δp^*_{t+1} denote domestic and foreign inflation rates over the same time period. Equation (6.5) is of central importance for this study. It summarizes the Backus-Smith condition for efficient international risk-sharing in the presence of real exchange rate changes: the difference between marginal utility growth rates for two countries should equal changes in their bilateral real exchange rate. In turn, it suggests that countries with appreciating currency (in real terms) should also experience relatively higher marginal utility growth.

Empirical Implementation: CRRA Utility

Equation (6.5) connects the movements of the bilateral real exchange rate with the difference between marginal utility growth rates for the corresponding countries. However, the latter are expressed through the general form of the utility function, which does not allow for direct empirical testing. Therefore, in order to empirically test this condition, a specific form has to be assigned to the utility function. For example, the utility function can belong to the general class of constant relative risk aversion (CRRA) functions, given by the following equation:

$$u(c_t) = \frac{(c_t)^{1-\gamma}}{1-\gamma} \tag{6.6}$$

where γ is the coefficient of risk aversion.[6] Using this form for the utility function, condition (6.4) can be rewritten as follows:

$$\frac{s_{t+1}}{s_t}\frac{P^*_{t+1}}{P^*_t}\frac{P_t}{P_{t+1}}\frac{(c_{t+1}^{-\gamma})}{(c_t^{-\gamma})}=\frac{(c_{t+1}^{*-\gamma})}{(c_t^{*-\gamma})} \tag{6.7}$$

or, in terms of rates of change:

$$\Delta(c_{t+1})-\Delta(c^*_{t+1})=\frac{1}{\gamma}(\Delta s_{t+1}+\Delta p^*_{t+1}-\Delta p_{t+1}) \tag{6.8}$$

Now, we define the difference between domestic and foreign levels of variable x by a tilde ($\tilde{x}$). Therefore, we get the following definitions $\Delta\tilde{c}_{t+1}\equiv\Delta(c_{t+1})-\Delta(c^*_{t+1})$ and $\Delta\tilde{p}_{t+1}\equiv\Delta(p_{t+1})-\Delta(p^*_{t+1})$.

Throughout subsequent empirical analysis, we use $\Delta\hat{p}_{t+1}\equiv\Delta(p^*_{t+1})-\Delta(p_{t+1})=-\Delta\tilde{p}_{t+1}$ rather than $\Delta\tilde{p}_{t+1}$. Using this notation, both components of the real exchange rate are positively related to relative consumption growth. As a result, we can simplify condition (6.8) as follows:

$$\Delta\tilde{c}_{t+1}=\frac{1}{\gamma}(\Delta s_{t+1}+\Delta\hat{p}_{t+1}) \tag{6.9}$$

In this way, equation (6.9) relates the relative consumption growth with the two components of the real exchange rate changes: the nominal exchange rate changes Δs_{t+1} and the negative of the inflation differential $\Delta\hat{p}_{t+1}$. Furthermore, the sum of the right hand side terms equals the change in the real exchange rate, so that the following equation holds:

$$\Delta\tilde{c}_{t+1}=\frac{1}{\gamma}(\Delta e_{t+1}) \tag{6.10}$$

Condition (6.10) implies a direct, testable relation between relative consumption growth and total real exchange rate changes. In fact, it gives a testable equation of the Backus-Smith condition: in the presence of real exchange rate fluctuations, efficient international risk-sharing implies that consumption growth should be higher in countries that experience relative drop in the price of consumption. Therefore, the major part of the empirical analysis in this chapter is based on (modified versions of) this equation.

[6] For the purpose of simplicity and tractability in the rest of the analysis, we assume that both countries have the same coefficient of risk aversion.

Empirical Specification(s)

Condition (6.10) implies the following general form of the regression equation to be estimated:

$$\Delta\tilde{c}_{it} = \beta_0 + \beta_1 \Delta e_{it} + u_{it} \tag{6.11}$$

where $\Delta\tilde{c}_{it}$ refers to the relative consumption growth (difference between domestic and foreign consumption growth rates) for country-pair i at time t, Δe_{it} refers to the real exchange real (price of foreign in terms of domestic goods) for country pair i at time t, and u_{it} is the error term (it is assumed that $u_{it} \sim iid(0, \sigma^2_{u_{it}})$).[7]

Hence, this general form of the regression equation gives two coefficients to be estimated: β_0 and β_1. The former gives the regression intercept,[8] while the latter gives the slope coefficient estimate. In turn, this slope coefficient is the main object of interest for this study because it measures the co-movement between real exchange rate changes and relative consumption growth. According to the theoretical framework with CRRA utility function elaborated above, this slope coefficient corresponds to the reciprocal value of the coefficient of relative risk-aversion. Therefore, β_1 is expected to be positive, significantly different from zero and equal to $1/\gamma$.

As mentioned above, equation (6.11) gives the very general empirical specification for testing the Backus-Smith condition. In fact, several modified versions of this general equation are used in the empirical analysis. Moreover, these modifications are based on three dimensions: first, with respect to the exchange rate regime (fixed vs. floating/flexible); second, with respect to different decompositions of the real exchange rate changes; and third, allowing for the possibility of partial or incomplete international risk-sharing.

[7] In general, this term should capture all non-consumption preference shocks, measurement errors, and relevant idiosyncratic characteristics of country-pair i not captured by the real exchange rate changes. Here we assume that these preference shocks and (consumption) measurement errors are not correlated across country-pairs.

[8] Strictly speaking, the general form of the regression equation does not include a constant term or intercept (see also Hess and Shin, 2006). However, we include it here in order to avoid possible technical problems. The estimation results are robust to the inclusion/exclusion of this constant term.

The first distinction concerns the differences that exist among regression equations estimated for alternative samples. For example, sample (1) refers to the Eurozone 12 countries after the introduction of the Euro. Since nominal exchange rates are constant for all countries in this sample, we estimate the following modified version of equation (6.11):

$$\Delta\tilde{c}_{it} = \beta_0 + \beta_1 \Delta\hat{p}_{it} + u_{it} \tag{6.12}$$

Second, for the other four samples characterized by flexible/floating nominal exchange rates, we use several alternative decompositions of the real exchange rate term. Initially, we include each component of the real exchange rate separately. Hence, we estimate the following two specifications:

$$\Delta\tilde{c}_{it} = \beta_0 + \beta_1 \Delta s_{it} + u_{it} \tag{6.13}$$

and

$$\Delta\tilde{c}_{it} = \beta_0 + \beta_1 \Delta\hat{p}_{it} + u_{it} \tag{6.14}$$

In this way, we investigate the individual effect of each real exchange rate component on the consumption growth differential. Furthermore, we include the total real exchange rate movement as the sole explanatory variable. Then, we estimate an equation similar to (6.11). Finally, we decompose the real exchange rate change into its two components: the nominal exchange rate change and the inflation rate differential. In this case, we investigate the individual contribution of each real exchange rate component by estimating the following type of regressions:

$$\Delta\tilde{c}_{it} = \beta_0 + \beta_1 \Delta s_{it} + \beta_2 \Delta\hat{p}_{it} + u_{it} \tag{6.15}$$

It is important to note that the efficient risk-sharing condition gives very clear implications for all slope coefficients in the alternative specifications (see equations (6.9) and (6.10)). According to the theoretical condition, the slope coefficient estimates in front of the inflation differential and the nominal exchange rate change should be very similar to the one in front of the total real exchange rate change term. Hence, they should be all positive and significantly different from zero.

Finally, the third modification of the general empirical specification concerns the possibility for partial risk-sharing. Various frictions in goods markets (non-tradeable goods, transportation costs, tariff and quota barriers,etc.) or asset markets (capital account restrictions, capital market imperfections, etc.) constrain the portion of aggregate, macroeconomic risks that can be potentially shared across borders. Therefore, we allow for incomplete or partial international risk-sharing in the general framework by including a relative output growth term as an additional explanatory variable. In fact, this term represents difference between the output growth rates for the corresponding countries and is defined in a similar way as the consumption growth differential:

$$\Delta\tilde{g}_{t+1} \equiv \Delta g_{t+1} - \Delta g^{*}_{t+1} \tag{6.16}$$

where $\Delta(g_{t+1})$ and $\Delta(g^{*}_{t+1})$ refer to domestic and foreign output growth rates, respectively. In this case, the general form of the regression equation testing the Backus-Smith condition looks as follows:

$$\Delta\tilde{c}_{it} = \beta_0 + \beta_1\Delta e_{it} + \beta_2\Delta\tilde{g}_{it} + u_{it} \tag{6.17}$$

If complete risk sharing takes place across countries, then the relative consumption growth should be unpredictable, and therefore uncorrelated with relative output growth. Therefore, in this case, the coefficient in front of the relative output growth term should not be significantly different from zero. On the other hand, a significantly positive estimate for this slope coefficient means that relative consumption responds and closely follows relative output growth. In turn, a significantly positive estimate for β_2 implies a deviation from complete international risk-sharing. The interpretation of the other slope coefficients in this regression equation is similar to the one given for the case of complete risk-sharing.

6.3 Data and Empirical Strategy

6.3.1 Data Sources

The dataset consists of quarterly observations for all 12 Eurozone countries as well as six major industrial countries outside the Eurozone: Australia, Japan, Sweden, Switzerland, UK, and USA. It includes data series on four different economic categories: real per capita consumption, real per capita gross domestic product, nominal exchange rates and inflation rates. Data on consumption and gross domestic product comes from the OECD Main Economic Indicators database. For consumption, we use the growth rate of real per capita final private consumption expenditure or the growth rate of real per capita personal expenditure, depending on the exact definition for the corresponding country. we calculate it as the quarterly rate of change in real (aggregate) final consumption expenditure deflated by the population growth rate. The growth rate of real per capital gross domestic product is calculated in a similar way. Data on nominal exchange rates comes from the GTIS database. For all countries in the dataset, we calculate quarterly changes in the nominal exchange rates (calculated through the US/domestic currency GTIS exchange rate series). The inflation rate series are calculated as quarterly changes in the general CPIs (alternatively NADJ). All data is retrieved from Datastream.

6.3.2 Data Samples

This study is primarily interested in empirically testing several relations for the 12 countries of the Eurozone during the Euro period. Nonetheless, in order to captire the specific nature of these relationships for the Eurozone, we compare it with four other samples (groups of countries/time periods). All of these samples are shown in Table 6.1. Sample (1) refers to the sample of primary interest, i.e. the countries of the Eurozone during the Euro period (1999-2006). First, the results for this sample are compared with the results for the same group of countries in different time periods. Thus, sample (2) covers a long period (1986-1998) and sample (3) covers a shorter period before the introduction of the Euro (1995-1998). This specific division of the entire time period was dictated by data availability issues. The former sample covers the

Table 6.1: Samples and Time Periods

Sample	Countries	Time period
(1)	Eurozone	1999:Q1-2006:Q1
(2)	Eurozone countries	1986:Q1-1998:Q4
(3)	Eurozone countries	1995:Q1-1998:Q4
(4)	Industrial countries	1986:Q1-2006:Q1
(5)	Industrial countries	1999:Q1-2006:Q1

longest period for which data was available, but this panel dataset is strongly unbalanced. The latter sample covers a much shorter period with a strongly balanced panel dataset. Second, the results from sample (1) are compared with two samples for a group of (major) industrial countries: sample (4) covering the entire time period (1986-2006), and sample (5) covering the Euro period only (1999-2006). Both data samples for the group of industrial countries are strongly balanced.[9]

6.3.3 Summary Statistics

The summary statistics are presented in Tables 6.2-6.4. Table 6.2 displays (annualized) means and (annualized) standard deviations of the three data series used in the estimations for the Eurozone countries after the introduction of the Euro (sample 1).[10] There are several interesting findings in Table 6.2. First, and in sharp contrast to macroeconomic models that assume (perfect) international capital market integration, the figures in this table suggest that the countries from the Eurozone do not exploit risk-sharing opportunities. In fact, consumption per capita growth is much more variable than GDP per capita growth for each sample in this table, just opposite of what theory would suggest. Second, the annualized growth rates for consumption and GDP per capita are of very similar magnitude and typically equal about 3 percent per annum.

Table 6.2 contains summary statistics for the Eurozone before the intro-

[9]There are only several missing observations in the consumption/output series for Japan and Sweden.

[10]The last row contains the cross-country average values for the summary statistics.

Table 6.2: Summary Statistics: Eurozone Countries (1999 : $Q1$-2006 : $Q1$)

Country	**C**		**Y**		**P**	
	Sample 1: 1999:Q1-2006:Q1					
	mean	st.dev	mean	st.dev	mean	st.dev
AUS	1.57	1.21	1.87	1.44	1.91	1.06
BEL	1.71	2.13	2.10	2.13	2.08	1.62
FIN	3.23	1.74	3.27	2.95	1.47	1.76
FRA	2.62	1.32	1.97	1.76	1.79	1.29
GER	0.66	2.87	1.24	1.99	1.58	1.34
GRE	7.72	17.49	4.56	5.41	3.34	5.81
ITA	0.99	2.03	1.34	1.84	3.72	2.50
IRE	5.33	5.05	5.83	7.23	2.37	0.71
LUX	3.91	10.14	4.45	4.79	2.52	1.99
NL	1.06	3.37	1.81	1.92	2.28	1.73
POR	1.74	2.82	3.71	1.64	3.18	3.13
SPA	7.23	3.25	1.33	3.05	3.04	2.19
MEAN	3.15	4.45	2.79	3.01	2.44	2.09

Note: The table presents summary statistics (means and standard deviations) for the following series: real per capital consumption growth rate (C), real per capita GDP growth (Y), and inflation rate (measures as percentage change in the CPI (P) for each of the 12 countries from the Eurozone. The sample period is 1999 : $Q1$-2006 : $Q1$ (Sample 1). All statistics are calculated using quarterly observations. The last row contains average values for the corresponding statistic across all countries. Quarterly data on real per capita GDP growth for Greece was not available over the entire time-period (they start only in 2000:Q1). All figures presented in the table are annualized and expressed in percentage terms (rounded to two decimal places).

duction of the Euro. It is divided into 2 panels: the upper one corresponds to sample (2), while the lower one corresponds to sample (3) from Table 6.1. There are broad similarities with the figures in Table 6.2: consumption per capita growth was similar in magnitude, but much more variable than GDP per capita growth. Moreover, inflation rates were much higher and more variable in the more distant past (upper panel) compared to the last decade (lower panel of Table 6.3 and Table 6.2). This observation at least partially reflects the general improvement in macroeconomic policy management in more recent past. Finally, although average changes in nominal exchange rates across the countries in this dataset were close to zero, they were also by far the most variable of all data series.[11] In fact, the average (annualized) standard deviation of the nominal exchange rate changes was about 24 percent for the entire time period, much higher than any other economic series.

Summary statistics for the group of industrial countries are reported in Table 6.4. Panel A contains figures for sample (4) covering the long period (1986 : $Q1$-2006 : $Q1$), while panel B contains figures for sample (5), covering the period after the introduction of the Euro (1999 : $Q1$-2006 : $Q1$).

[11] All nominal exchange rate changes are calculated against the US dollar.

Table 6.3: Summary Statistics: Eurozone Countries

Country	C		Y		P		S	
	Sample 2: 1986:Q1-1998:Q4							
	mean	st.dev	mean	st.dev	mean	st.dev	mean	st.dev
AUS	2.37	1.37	2.80	1.34	2.34	2.97	3.43	25.74
BEL	2.00	1.74	2.20	1.77	2.08	1.59	3.29	24.91
FIN	1.86	3.91	2.26	5.16	2.85	2.71	0.80	22.70
FRA	1.67	2.33	2.06	1.85	2.19	1.30	2.66	23.86
GER	1.85	4.86	1.38	2.96	2.29	2.44	3.48	25.76
GRE	2.40	5.95			11.56	8.55	-4.94	20.96
ITA	2.16	2.94	10.04	11.73	2.41	1.78	1.69	22.85
IRE	8.12	8.58	1.88	2.25	4.40	1.92	0.42	24.40
LUX	4.08	10.53	4.77	5.74	0.89	0.99	3.29	24.91
NL	4.43	1.92	3.92	1.80	1.90	1.89	3.42	25.52
POR	4.23	3.78	3.31	1.95	4.56	2.46	0.92	23.88
SPA	5.96	3.87	4.00	1.87	6.93	4.65	-0.42	21.87
MEAN	3.43	4.32	3.51	3.49	3.70	2.77	1.50	23.95
	Sample 3: 1995:Q1-1998:Q4							
	mean	st.dev	mean	st.dev	mean	st.dev	mean	st.dev
AUS	1.58	1.12	2.79	1.01	1.47	2.16	-1.19	19.74
BEL	1.99	1.74	2.20	1.77	1.46	1.65	-1.39	19.95
FIN	3.70	1.17	4.69	2.04	0.82	1.14	-0.91	18.34
FRA	1.86	3.02	1.99	1.33	1.19	1.21	-0.90	18.32
GER	1.70	2.47	1.54	2.58	1.31	1.66	-1.22	19.99
GRE	2.40	5.95			5.61	6.50	-4.05	16.98
ITA	2.23	2.67	1.43	2.17	1.80	1.77	-0.82	15.19
IRE	8.12	8.58	10.04	11.73	2.93	1.93	-0.92	17.31
LUX	4.098	10.53	4.77	5.74	0.89	0.99	-1.39	19.95
NL	4.43	1.92	3.91	1.79	1.98	0.92	-1.38	19.89
POR	4.23	3.78	3.30	1.95	2.76	1.79	-1.99	16.62
SPA	5.96	3.87	4.01	1.87	2.94	2.27	-1.46	17.79
MEAN	3.52	3.90	3.39	2.83	2.10	2.01	-1.47	18.34

Note: The table presents summary statistics (means and standard deviations) for the following series: real per capital consumption growth rate (C), real per capita GDP growth (Y), inflation rate (measures as percentage change in the CPI (P), and nominal exchange rate changes (S) for each of the 12 countries from the Eurozone. All statistics are calculated using quarterly observations. The nominal exchange rate changes are calculated with respect to the US dollar (units of currency per US dollar). Each panel refers to a different sample period: the upper panel refers to period $1986 : Q1$-$1998 : Q4$ (Sample 2), while the lower panel refer to period $1995 : Q1$-$1998 : Q4$ (Sample 3). Consumption and GDP growth series for most countries start only in 1995. Therefore, most of the summary statistics for these series are same in both panels. The last row in each panel contains average values for the corresponding statistic across all countries. Quarterly data on real per capita GDP growth for Greece was not available over the entire time-period (they start only in 2000:Q1), hence the empty spaces in panels B and C. All figures presented in the table are annualized and expressed in percentage terms (rounded to two decimal places).

In contrast to the previous table, per capita consumption growth shows similar levels of variability compared to per capita GDP growth. Moreover, per capita consumption growth variability is slightly lower than per capita GDP growth variability in the latter period. In turn, this suggest better risk-sharing for the countries in this group. The other figures are comparable to those for the Eurozone. For example, inflation strongly decreased and became considerably less variable in the more recent past: the average inflation rate decreased from 2.57 to 1.7, and its standard deviation fell from 2.77 to 1.87

Table 6.4: Summary Statistics: Industrial Countries

Country	C		Y		P		S	
Panel A: 1986:Q1-2006:Q1								
	mean	st.dev	mean	st.dev	mean	st.dev	mean	st.dev
AUT	3.28	2.36	3.38	2.71	3.56	3.15	0.58	20.09
JAP	1.14	3.35	1.28	2.96	0.60	2.54	2.94	26.47
SWE	2.26	2.64	2.99	1.73	2.89	3.91	0.26	22.81
SWI	1.39	2.14	1.54	2.70	1.81	2.35	2.87	26.55
UK	2.97	2.74	2.55	1.98	3.50	3.31	1.28	20.06
USA	3.29	1.89	3.04	2.03	3.07	1.37		
MEAN	2.39	2.52	2.46	2.35	2.57	2.77	1.59	23.19
Panel B: 1999:Q1-2006:Q1								
	mean	st.dev	mean	st.dev	mean	st.dev	mean	st.dev
AUT	3.71	1.83	3.17	2.14	3.17	2.53	3.29	23.89
JAP	1.18	1.84	1.63	2.57	-0.42	1.28	0.06	22.62
SWE	2.39	2.74	2.89	1.81	1.31	2.10	0.85	22.05
SWI	1.48	1.51	1.81	2.63	0.98	2.13	1.21	22.47
UK	3.01	1.99	2.71	1.11	2.45	1.99	0.77	14.84
USA	3.41	1.67	2.79	2.15	2.74	1.22		
MEAN	2.53	1.93	2.50	2.07	1.70	1.87	1.24	21.17

Note: The table presents summary statistics (means and standard deviations) for the following series: real per capital consumption growth rate (C), real per capita GDP growth (Y), inflation rate (measures as percentage change in the CPI (P), and nominal exchange rate changes (S) for 6 industrial countries. All statistics are calculated using quarterly observations. The nominal exchange rate changes are calculated with respect to the US dollar (units of currency per US dollar). Panel A refers to the long time period 1986 : $Q1$-2006 : $Q1$, while Panel B refers to the period after the introduction of the Euro: 1999 : $Q1$-2006 : $Q1$. Complete consumption and GDP growth series over the entire time period are available for the following countries: Australia, Switzerland, UK, and USA. The corresponding series for Japan and Sweden start in 1994 : $Q1$ and 1993 : $Q1$, respectively. All figures presented in the table are annualized and expressed in percentage terms (rounded to two decimal places).

percent. Moreover, as in the case of the Eurozone, nominal exchange rates were by far the most variable of all time series.

6.4 Results

This section presents the main findings from the empirical analysis. It is divided into two parts: the first one presents regression results about the Eurozone countries in the Euro-period, and the second one presents results from similar empirical specifications for alternative data samples (Eurozone countries in the pre-Euro-period, and several other industrial countries).

6.4.1 Eurozone Countries in the Euro-Period

Bilateral Estimations

In order to investigate the relation between relative consumption growth and real exchange rate changes, we estimate bilateral regressions for each country of the Eurozone against all other 11 partner countries. The results from these estimations are presented in Table 6.5. For each pair of countries, given by the column-row intersection, we present the slope coefficient estimates and its corresponding t-statistic (in brackets). Therefore, these figures show the severity of the Backus-Smith puzzle for the Eurozone countries: significantly negative coefficients indicate presence of the Backus-Smith puzzle, while significantly positive ones support the theory.

There are mixed results in this table. In fact, the coefficient estimates are still (significantly) negative or insignificantly different from zero for several bilateral country pairs, as suggested by the large literature on the Backus-Smith puzzle. Nonetheless, the majority of coefficient estimates are positive, and many of them are significantly different from zero, just in accordance with theory.

In fact, most of the significantly negative results refer to the bilateral pairs that include Luxembourg. If this country is taken aside, then there is only one significantly negative coefficient estimate in the entire dataset: the one corresponding to the country pair Spain-Portugal. Contrary to this, there are over a dozen of positive coefficient estimates significantly different from zero. These results clearly give support for the theory and suggest that the Backus-Smith anomaly has largely disappeared for the Eurozone countries after the introduction of the common currency. In order to further investigate this suggestive finding, we present estimation results for the complete panel dataset.

Table 6.5: Results from Bilateral Estimations: Eurozone 1999:Q1-2006:Q1

	AUS	**BEL**	**FIN**	**FRA**	**GER**	**GRE**	**IRE**	**ITA**	**LUX**	**NL**	**POR**	**SPA**
AUS												
BEL	0.46											
	(2.09)**											
FIN	0.346	0.391										
	(1.33)	(1.26)										
FRA	0.316	0.347	0.378									
	(1.33)	(1.00)	(1.33)									
GER	0.732	1.296	0.5	0.312								
	(1.74)*	(2.60)**	(1.76)*	(0.93)								
GRE	1.036	0.969	1.369	1.017	0.861							
	(1.83)*	(1.82)*	(2.16)**	(1.65)	(1.62)							
IRE	0.445	0.714	0.258	-0.697	1.328	0.683						
	(0.46)	(1.05)	(0.38)	(-0.80)	(1.64)	(1.25)						
ITA	-0.183	0.03	-0.14	-0.219	0.172	1.224	-0.275					
	(-1.09)	(0.15)	(-0.42)	(-1.02)	(0.70)	(1.99)*	(-0.66)					
LUX	-3.633	-3.277	-2.489	-4.028	-2.826	2.928	-4.209	-1.082				
	(-5.08)***	(-5.93)***	(-2.32)**	(-4.18)***	(-5.26)***	(3.32)***	(-5.04)***	(-1.03)				
NL	-0.272	-0.062	0.456	-0.165	0.42	0.758	-0.885	-0.184	-2.439			
	(-0.69)	(-0.15)	(0.91)	(-0.39)	(0.80)	(1.34)	(-1.06)	(-0.57)	(-3.04)***			
POR	0.016	0.042	0.086	-0.092	0.086	0.76	-0.382	-0.265	0.225	-0.086		
	(0.09)	(0.24)	(0.36)	(-0.42)	(0.36)	(1.19)	(-1.11)	(-1.19)	(0.27)	(-0.33)		
SPA	0.625	0.424	0.928	0.586	0.825	0.539	0.253	-0.467	0.171	0.376	-0.776	
	(2.66)***	(1.93)*	(2.87)***	(1.90)*	(3.15)***	(0.72)	(0.46)	(-1.21)	(0.11)	(0.78)	(-2.15)**	

Note: The table presents results from bilateral time-series estimations given by the following regression equation $\Delta\tilde{c}_t = \beta_0 + \beta_1\Delta\hat{p}_t + u_t$ among all 12 Eurozone countries. All results are based on 28 quarterly observations over the period 1999:Q1-2006:Q1. For each pair of countries, given by the column-row intersection, we present the slope coefficient estimates and its corresponding t-statistic (in brackets). Significance at 10%, 5%, and 1% is indicated by *, **, and ***, respectively. The critical values for the t-distribution with 27 degrees of freedom are 1.703 (10%), 2.052 (5%), and 2.771 (1%).

Pooled-OLS and Panel Evidence

Table 6.6 presents results from pooled-OLS and panel estimations for the Eurozone countries in the period after the introduction of the Euro.[12] It is divided into three panels, each corresponding to a different set of countries. The columns contain results for different estimation procedures: pooled-OLS, panel estimation with random effects and panel estimation with fixed effects.[13] Moreover, for each of these specifications, we allow for partial (incomplete) risk-sharing (columns marked by (g)).[14]

Panel A refers to the complete dataset - all 12 Eurozone countries. The first row in this panel gives coefficient estimates for the real exchange rate, i.e. the inflation differential in this time period. They are positive and significantly different from zero for each specification that does not allow for partial risk-sharing. In fact, the t-statistics (given in brackets) are far above the usual critical values, suggesting that this relation is statistically significant at any conventional level (1 percent, for example). When we allow for incomplete risk-sharing, the slope coefficient estimates for the output term are significantly positive in each of the three specifications, suggesting that international risk-sharing in this country set is far from perfect.[15] Simultaneously, the slope coefficient in front of the real exchange rate (the negative of the inflation differential) remains positive and statistically significant for all but one specification (pooled-OLS).[16] This is the most important result

[12] We only report results from pooled-OLS estimations for sample (1) and not for the other samples. This estimation procedure usually produces biased coefficient estimates when one does not control for the country-pair specific portion of the error term. The results from the pooled-OLS estimations for the remaining samples are available upon request.

[13] The fixed-effects in these estimations refer to the country-pair-specific part of the error term. The country-pairs are basic units here, because the analysis is based on pairwise comparisons.

[14] Strictly speaking, the coefficients in front of the output growth terms Δg_{t+1} and Δg^*_{t+1} might differ and should not be restricted to equal each other. Relaxing this restriction, however, does not lead to (qualitative) changes in the main results.

[15] The adjustment of the standard errors for intragroup (intertemporal) serial correlation produces somewhat lower t-values for the coefficient in front of the relative output growth terms, but does not lead to qualitatively different conclusions.

[16] Pooled-OLS estimation are usually inconsistent in this setting. They are presented for illustrative purposes here.

in this study: when nominal exchange rates are fixed, relative consumption co-moves positively with the real exchange rate, i.e. the Backus-Smith puzzle disappears.

The lower two panels restrict the complete dataset: the results in Panel B exclude Luxembourg, while those in Panel C exclude Luxembourg and Greece. These exclusions are based on several arguments about the countries concerned. First, the results from the bilateral estimations presented in Table 6.5 already suggested that Luxembourg behaves very differently from the rest of the Eurozone countries. Furthermore, there are several reasons that might justify the exclusion of Luxembourg from this (pooled) analysis: it is much smaller (geographically and economically) than the other countries, has a very high concentration of international institutions and international banks, and employs a very mobile workforce.[17] Second, there are several limitations in the quarterly data series for Greece used in the analysis, motivating its exclusion in panel C.[18] Moreover, Table 6.3 suggests that most of these macroeconomic series display persistently higher levels and/or inflation compared to most other countries even into the Euro-period.

Several findings in these lower panels deserve attention. First, the main result from Panel A does not vanish. In fact, the slope coefficient estimate in front of the real exchange rate retains its positive sign and statistical significance in all specifications. Moreover, its economic and statistical significance increases in the second panel, while it drops only marginally in the third panel. This indicates that the main finding from Panel A is not due to the inclusion of one of these countries. Second, the coefficient on the relative output drops marginally in the second panel, while it completely loses its significance in the last panel. Therefore, this finding suggests that once the effect of the real exchange rate is taken into account, the hypothesis of perfect risk-sharing for this restricted dataset cannot be rejected.

In order to deal with possible endogeneity issues, we instrument the inde-

[17] Many of these aspects can help in explaining why the macroeconomic variables for Luxembourg might not correctly reflect the changes in purchasing power of its inhabitants as well as the correlation between relative purchasing power changes and relative consumption growth.

[18] The quarterly consumption growth series for Greece are available only from 1995:Q1, while the quarterly real GDP growth series are available only from 2000:Q1.

Table 6.6: Pooled OLS and Panel Estimations: Eurozone 1999:Q1-2006:Q1

	pooled	pooled (g)	panel re	panel fe	panel re (g)	panel fe (g)
			Panel A: All country pairs			
$\hat{p}$	0.251	0.062	0.385	0.437	0.115	0.126
	(4.93)***	(1.71)*	(7.65)***	(8.59)***	(3.28)***	(3.54)***
$\tilde{g}$		0.107			0.145	0.153
		(5.02)***			(6.78)***	(6.97)***
Obs	3674	3586	3674	3674	3586	3586
R^2	0.01	0.01		0.02		0.02
Stacks			132	132	132	132
			Panel B: Excluding Luxembourg			
$\hat{p}$	0.382	0.21	0.548	0.592	0.306	0.315
	(7.73)***	(7.07)***	(11.47)***	(12.30)***	(11.82)***	(12.16)***
$\tilde{g}$		0.037			0.071	0.073
		(1.98)**			(4.24)***	(4.34)***
Obs	3080	3000	3080	3080	3000	3000
R^2	0.02	0.02		0.05		0.07
Stacks			110	110	110	110
			Panel C: Excluding Luxembourg and Greece			
$\hat{p}$	0.151	0.083	0.263	0.277	0.18	0.188
	(3.26)***	(2.23)**	(6.05)***	(6.34)***	(5.49)***	(5.72)***
$\tilde{g}$		0.079			0.004	0.013
		(3.88)***			(0.24)	(0.70)
Obs	2660	2640	2660	2660	2640	2640
R^2	0	0.01		0.02		0.01
Stacks			95	95	95	95

Note: The table presents results from pooled-OLS and panel data estimations (with random-effects and fixed-effects) for the Eurozone countries in the period after the introduction of the Euro: 1999:Q1-2006:Q1. The general regression specification is given by the following equation: $\Delta\tilde{c}_{it} = \beta_0 + \beta_1\Delta\hat{p}_{it} + u_{it}$, while the specification that allows for partial risk-sharing in consumption is given by: $\Delta\tilde{c}_{it} = \beta_0 + \beta_1\Delta\hat{p}_{it} + \beta_2\Delta\tilde{g}_{it} + u_{it}$. Panel A contains results for all 12 countries, Panel B excludes Luxembourg, and Panel C excludes Luxembourg and Greece. The number of stacks refers to the number of bilateral country-pairs included in the panel. For each specification, the table displays the slope coefficient estimates with the corresponding t-statistics (in brackets). Significance at 10%, 5%, and 1% is indicated by *, **, and ***, respectively. The critical values for the t-distribution are 1.645 (10%), 1.960 (5%), and 2.576 (1%).

pendent variables with the lagged value of the dependent variable ($\Delta\tilde{c}_{i,t-1}$) and the lagged values of all independent variables ($\Delta\hat{p}_{i,t-1}$ and $\Delta\tilde{g}_{i,t-1}$ in this case).[19]

Table 6.7 presents results from the instrumental variables panel estimations. In accordance with the set-up of Table 6.6, it is divided into three panels: complete set of Eurozone countries (Panel A), excluding Luxembourg

[19]For similar choice of instruments see Hess and Shin (2006). The results from the first-stage regression suggest strong correlation between the variables chosen as instruments and the original explanatory variables. For example, the value for the F-statistic from the first-stage regression of $\Delta\hat{p}_{it}$ on the three instruments is 327.85, while the corresponding value for the first-stage regression of $\Delta\tilde{g}_{it}$ on the same three instruments is 181.62. Both values are much higher than 10, a threshold value required for strong instruments. All results are available upon request.

Table 6.7: Instrumental Variables (IV) Panel Estimations: Eurozone 1999:Q1-2006:Q1

	panel re	panel fe	panel re (g)	panel fe (g)
		Panel A: All country pairs		
$\hat{p}$	0.39	0.442	0.122	0.133
	(7.72)***	(8.64)***	(3.44)***	(3.68)***
$\tilde{g}$			0.141	0.148
			(6.51)***	(6.71)***
Obs	3651	3651	3586	3586
R^2		0.02		0.02
Stacks	132	132	132	132
		Panel B: Excluding Luxembourg		
$\hat{p}$	0.549	0.593	0.312	0.321
	(11.47)***	(12.29)***	(11.98)***	(12.29)***
$\tilde{g}$			0.069	0.071
			(4.09)***	(4.19)***
Obs	3067	3067	2968	2968
R^2		0.05		0.07
Stacks	110	110	110	110
		Panel C: Excluding Luxembourg and Greece		
$\hat{p}$	0.265	0.28	0.185	0.193
	(6.11)***	(6.39)***	(5.64)***	(5.87)***
$\tilde{g}$			0.003	0.011
			(0.16)	(0.61)
Obs	2648	2648	2623	2623
R^2		0.02		0.01
Stacks	95	95	95	95

Note: The table presents results from instrumental variables panel data estimations (with random-effects and fixed-effects) for the Eurozone countries in the period after the introduction of the Euro: 1999:Q1-2006:Q1. The lagged values of the dependent variable $\Delta\tilde{c}_{i,t-1}$ and/or the lagged values of the independent variables $\Delta\hat{p}_{i,t-1}$ and $\Delta\tilde{g}_{i,t-1}$ are used as instruments for the original independent variables. The general regression specification is given by the following equation: $\Delta\tilde{c}_{it} = \beta_0 + \beta_1\Delta\hat{p}_{it} + u_{it}$, while the specification that allows for partial risk-sharing in consumption is given by: $\Delta\tilde{c}_{it} = \beta_0 + \beta_1\Delta\hat{p}_{it} + \beta_2\Delta\tilde{g}_{it} + u_{it}$. Panel A contains results for all 12 countries, Panel B excludes Luxembourg, and Panel C excludes Luxembourg and Greece. The number of stacks refers to the number of bilateral country-pairs included in the panel. For each specification, the table displays the slope coefficient estimates with the corresponding t-statistics (in brackets). Significance at 10%, 5%, and 1% is indicated by *, **, and ***, respectively. The critical values for the t-distribution are 1.645 (10%), 1.960 (5%), and 2.576 (1%).

(Panel B), and excluding Greece and Luxembourg (Panel C). The main conclusions from Table 6.6 stay unchanged: relative consumption growth is positively related to the real exchange rate, and there is a significant departure from perfect risk-sharing for all but the last set of countries (Panel C).

6.4.2 Comparison with Alternative Samples

The last section presented results for the countries of the Eurozone after the introduction of the Euro. They clearly indicated that the Backus-Smith puzzle disappears in this sample. In order to see whether this result is due to the elimination of the nominal exchange rate, we contrast the findings for the

Eurozone in the period 1999:Q1-2006:Q1 with four alternative samples given in Table 6.1. First, we present results for the same set of countries (Eurozone 12) during alternative time periods (samples (2) and (3)), and second, we present evidence on a different set of industrial countries during the same time period (samples (4) and (5)).

Eurozone (1986:Q1-1998:Q4)

Table 6.8 presents estimation results for the same Eurozone countries during the entire time period before the introduction of the Euro (1986:Q1-1998:Q4).[20] All figures in this table refer to the fixed-effects panel estimations.[21] In accordance with the set-up for the first sample, the table is divided into three panels.

Each column of Table 6.8 corresponds to a different fixed-effects panel specification.[22] We regress the relative (bilateral) consumption growth on four sets of explanatory variables: i) (negative of the) inflation differential, ii) nominal exchange rate changes, iii) real exchange rate changes, and iv) joint inclusion of both components of real exchange rate changes - the inflation differential and the nominal exchange rate changes. Moreover, we estimate each of these four specifications by assuming perfect risk-sharing as well as by allowing for partial risk-sharing through the inclusion of the relative output growth term.[23]

There are several interesting findings in Table 6.8. Most importantly, the

[20] The specification(s) that exclude the nominal exchange rate have been estimated over the period 1970:Q1-1998:Q4 as well. There are no significant differences in the estimation results with respect to the exact starting date (1970:Q1 or 1986:Q1).

[21] We only report results from fixed-effects panel estimations. The fixed-effects refer to the country pair-specific, time-invariant components of the error term. In most specifications, the Hausman test suggests rejection of the null hypothesis of no systematic difference between estimates obtained with fixed-effects and estimates obtained with random-effects. Therefore, only fixed-effects estimations yield consistent estimates in this case. In some specifications, the Hausman test is (marginally) not rejected, but then the results from the two estimation procedures do not differ significantly and do not lead to qualitatively different conclusion(s).

[22] We do not report results from the instrumental variables estimations here since they do not differ significantly from the standard fixed-effects panel estimates. They are available upon request.

[23] The specifications that include the relative output term are marked with (g) as before.

results in this table present evidence about the "dichotomy" that exists between the two components of the real exchange rate: the inflation differential on one and the nominal exchange rate changes on the other hand. In fact, the coefficient for the inflation differential is positive and significantly different from zero in all specifications in Panel A. Clearly, this "macroeconomic" part of the real exchange rate behaves in line with theory. On the other hand, the coefficient for the nominal (and real) exchange rate is negative and significantly different from zero in each specification. Therefore, these two findings clearly suggest that the anomalous negative correlation between relative consumption growth and real exchange rate changes (Backus-Smith puzzle) comes from the nominal exchange rate behavior.

The lower two panels in this table report results for the group of Eurozone countries excluding Luxembourg (Panel B) and excluding Luxembourg and Greece (Panel C). The results in these panels are very similar and convey the same message as Panel A.[24] In fact, the coefficient in front of the inflation differential stays positive in all specifications. Moreover, it is statistically significant (at 5 percent level) in all but one specification. The results about the nominal exchange rate stay literally unchanged: its slope coefficient is negative and statistically significant at any conventional level (1 percent significance).[25]

Eurozone(1995:Q1-1998:Q4)

Table 6.8 presented results for the Eurozone countries during the entire period 1986:Q1-1998:Q4. However, the panel dataset used in these estimations is strongly unbalanced, since the data series do not have equal length for each country. Therefore, we turn to a shorter time period for which a balanced panel can be constructed.[26] Table6.9 presents results for this period, which refers to the four years before the introduction of the Euro (1995:Q1-1998:Q4). The table follows a very similar set-up like the previous tables.

[24]This comes (partly) as a result of the data limitations for these countries.

[25]A comparison between the slope coefficient estimates in front of the relative output growth term (g) (panels B and C in tables 6.6 and 6.8) suggests that, as expected, risk-sharing across Eurozone countries increases in the latter, Euro-period.

[26]There is a necessary trade-off between the time length of the dataset and data availability. Hence, the time dimension of the balanced panel dataset analyzed in these tables is 16 quarters.

Table 6.8: Panel Estimations: Eurozone countries 1986:Q1-1998:Q4

				Panel A: All countries				
$\hat{p}$	0.127	0.199					0.107	0.178
	(3.18)***	(2.85)***					(2.68)***	(2.54)**
$\tilde{g}$		0.270		0.271		0.272		0.267
		(8.56)***		(8.57)***		(8.62)***		(8.47)***
s			-0.063	-0.05			-0.061	-0.045
			(-5.71)***	(-2.92)***			(-5.44)***	(-2.61)***
e					-0.052	-0.038		
					(-4.75)***	(-2.21)**		
Obs	2442	1734	2442	1734	2442	1734	2442	1734
R^2	0.01	0.05	0.01	0.05	0.01	0.05	0.02	0.05
Stacks	132	110	132	110	132	110	132	110
				Panel B: Excluding Luxembourg				
$\hat{p}$	0.058	0.148					0.036	0.121
	(1.73)*	(2.71)***					(1.07)	(2.22)**
$\tilde{g}$		0.246		0.241		0.244		0.237
		(9.04)***		(8.91)***		(9.03)***		(8.75)***
s			-0.073	-0.061			-0.073	-0.058
			(-7.93)***	(-4.71)***			(-7.81)***	(-4.44)***
e					-0.067	-0.053		
					(-7.32)***	(-4.04)***		
Obs	2106	1446	2106	1446	2106	1446	2106	1446
R^2	0	0.06	0.03	0.07	0.03	0.07	0.03	0.08
Stacks	110	90	110	90	110	90	110	90
				Panel C: Excluding Luxembourg and Greece				
$\hat{p}$	0.099	0.148					0.072	0.121
	(2.28)**	(2.71)***					(1.68)*	(2.22)**
$\tilde{g}$		0.246		0.241		0.244		0.237
		(9.04)***		(8.91)***		(9.03)***		(8.75)***
s			-0.062	-0.061			-0.06	-0.058
			(-5.94)***	(-4.71)***			(-5.73)***	(-4.44)***
e					-0.055	-0.053		
					(-5.35)***	(-4.04)***		
Obs	1874	1446	1874	1446	1874	1446	1874	1446
R^2	0	0.06	0.02	0.07	0.02	0.07	0.02	0.08
Stacks	95	90	95	90	95	90	95	90

Note: The table presents results from fixed-effects panel data estimations for the Eurozone countries over the sample period 1986:Q1-1998:Q4. The general specification is given by the following regression equation: $\Delta\tilde{c}_{it} = \beta_0 + \beta_1 \Delta x_{it} + u_{it}$, while the specification that allows for partial risk-sharing in consumption is given by: $\Delta\tilde{c}_{it} = \beta_0 + \beta_1 \Delta x_{it} + \beta_2 \Delta\tilde{g}_{it} + u_{it}$. Moreover, each of these specifications is estimated using alternative variables in place of Δx_{it}: the negative of the inflation differential $\Delta\hat{p}_{it}$ (first two columns), the change in the nominal exchange rate Δs_{it} (third and fourth column), the change in the real exchange rate Δe_{it} (fifth and sixth column), and its both components $\Delta\hat{p}_{it}$ and Δs_{it} together (last two columns). Panel A contains results for all 12 countries, Panel B excludes Luxembourg, and Panel C excludes Luxembourg and Greece. The number of stacks refers to the number of bilateral country-pairs included in the panel. For each specification, the table displays the slope coefficient estimates with the corresponding t-statistics (in brackets). Significance at 10%, 5%, and 1% is indicated by *, **, and ***, respectively. The critical values for the t-distribution are 1.645 (10%), 1.960 (5%), and 2.576 (1%).

Table 6.9: Panel Estimations: Eurozone countries 1995:Q1-1998:Q4

				Panel A: All countries				
$\hat{p}$	0.162	0.33					0.151	0.34
	(3.48)***	(3.59)***					(3.23)***	(3.68)***
$\tilde{g}$		0.195		0.204		0.203		0.195
		(5.45)***		(5.72)***		(5.69)***		(5.46)***
s			-0.043	0.025			-0.039	0.035
			(-2.97)***	(0.89)			(-2.68)***	(1.23)
e					-0.026	0.055		
					(-1.85)*	(1.96)*		
Obs	1956	1468	1956	1468	1956	1468	1956	1468
R^2	0.01	0.03	0	0.02	0	0.03	0.01	0.03
Stacks	132	110	132	110	132	110	132	110
				Panel B: Excluding Luxembourg				
$\hat{p}$	0.08	0.285					0.066	0.287
	(2.08)**	(4.00)***					(1.72)*	(4.02)***
$\tilde{g}$		0.129		0.144		0.145		0.13
		(4.29)***		(4.78)***		(4.82)***		(4.31)***
s			-0.057	0.004			-0.055	0.01
			(-4.72)***	(0.17)			(-4.57)***	(0.48)
e					-0.046	0.028		
					(-3.96)***	(1.34)		
Obs	1620	1180	1620	1180	1620	1180	1620	1180
R^2	0	0.03	0.01	0.02	0.01	0.02	0.02	0.03
Stacks	110	90	110	90	110	90	110	90
				Panel C: Excluding Luxembourg and Greece				
$\hat{p}$	0.176	0.285					0.172	0.287
	(3.22)***	(4.00)***					(3.12)***	(4.02)***
$\tilde{g}$		0.129		0.144		0.145		0.13
		(4.29)***		(4.78)***		(4.82)***		(4.31)***
s			-0.016	0.004			-0.011	0.01
			(-1.06)	(0.17)			(-0.73)	(0.48)
e					-0.002	0.028		
					(-0.16)	(1.34)		
Obs	1388	1180	1388	1180	1388	1180	1388	1180
R^2	0.01	0.03	0	0.02	0	0.02	0.01	0.03
Stacks	95	90	95	90	95	90	95	90

Note: The table presents results from fixed-effects panel data estimations for the Eurozone countries over the sample period 1995:Q1-1998:Q4. The very short time period (16 observations per country) is chosen in order to have a strongly balanced pre-Euro dataset. The general specification is given by the following regression equation: $\Delta\tilde{c}_{it} = \beta_0 + \beta_1\Delta x_{it} + u_{it}$, while the specification that allows for partial risk-sharing in consumption is given by: $\Delta\tilde{c}_{it} = \beta_0 + \beta_1\Delta x_{it} + \beta_2\Delta\tilde{g}_{it} + u_{it}$. Moreover, each of these specifications is estimated using alternative variables in place of Δx_{it}: the negative of the inflation differential $\Delta\hat{p}_{it}$ (first two columns), the change in the nominal exchange rate Δs_{it} (third and fourth column), the change in the real exchange rate Δe_{it} (fifth and sixth column), and its both components $\Delta\hat{p}_{it}$ and Δs_{it} together (last two columns). Panel A contains results for all 12 countries, Panel B excludes Luxembourg, and Panel C excludes Luxembourg and Greece. The number of stacks refers to the number of bilateral country-pairs included in the panel. For each specification, the table displays the slope coefficient estimates with the corresponding t-statistics (in brackets). Significance at 10%, 5%, and 1% is indicated by *, **, and ***, respectively. The critical values for the t-distribution are 1.645 (10%), 1.960 (5%), and 2.576 (1%).

The panel estimation results in Table 6.9 strengthen the evidence about the "dichotomy" that exists between the two components of the real exchange rate. In fact, the inflation differential enters with a significantly positive sign in all specifications (always significantly different from zero at least at 10 percent level), while the nominal exchange rate enters with a significantly negative sign in all specifications that do not allow for partial risk-sharing (i.e. do not include the relative output growth term). Moreover, the coefficient estimates for the real exchange rate are very similar to those for the nominal exchange rate, supporting the proposition that the consumption real exchange rate anomaly is primarily due to the nominal exchange rate.

There are two main conclusions that can be drawn from Table 6.9. First, the results support the proposition that the nominal exchange rate is the main source of the consumption real exchange rate correlation (Backus-Smith) puzzle. Second, similar to the results for the Eurozone countries over the entire time period (sample 2), the results suggest that the negative of the inflation differential exhibited positive correlation with the relative consumption growth term, just as expected by macroeconomic theory (RBC models).

Industrial Countries

The results so far concerned only one group of countries (Eurozone 12) over three different time periods. The purpose was to show the importance of the nominal exchange rate changes for the consumption-real exchange rate (Backus-Smith) puzzle. Therefore, it was natural to start with a comparison of different episodes for the same group of countries: the Euro-period (sample (1)) was compared with two pre-Euro periods (samples (2) and (3)). Here, we do a similar exercise and compare the results for the Eurozone after the introduction of the Euro (sample (1)) with the results for a group of advanced industrial countries over the same Euro-period, and over a longer time period (1986:Q1-2006:Q1). We include six industrial countries that generally had flexible nominal exchange rates over the concerned time period: Australia, Japan, Sweden, Switzerland, UK and USA.

First, we start with an analysis of the entire time period over which data

Table 6.10: Industrial Countries 1986:Q1-2006:Q1

	Panel Estimates (Fixed Effects)							
$\hat{p}$	0.149	0.114					0.15	0.113
	(5.60)***	(4.61)***					(5.62)***	(4.52)***
$\tilde{g}$		0.411		0.422		0.422		0.412
		(16.35)***		(16.74)***		(16.70)***		(16.36)***
s			0.001	-0.003			-0.001	-0.002
			(0.09)	(-1.07)			(-0.42)	(-0.64)
e					0.002	-0.002		
					(0.59)	(-0.52)		
Obs	1844	1844	1844	1844	1844	1844	1844	1844
R^2	0.02	0.14	0	0.13	0	0.13	0.02	0.14
Stacks	30	30	30	30	30	30	30	30

Note: The table presents results from fixed-effects panel data estimations for 6 industrial countries over the sample period 1986:Q1-2006:Q1. The dataset is strongly unbalanced. Observations over the entire time period (1986:Q1-2006:Q1) are available for the following countries only: Australia, Switzerland, UK and USA, while for Japan and Sweden observations are available over the period 1994 : Q1-2006:Q1 and 1993 : Q1-2006:Q1, respectively. The general specification is given by the following regression equation: $\Delta\tilde{c}_{it} = \beta_0 + \beta_1 \Delta x_{it} + u_{it}$, while the specification that allows for partial risk-sharing in consumption is given by: $\Delta\tilde{c}_{it} = \beta_0 + \beta_1 \Delta x_{it} + \beta_2 \Delta\tilde{g}_{it} + u_{it}$. Moreover, each of these specifications is estimated using alternative variables in place of Δx_{it}: the negative of the inflation differential $\Delta\hat{p}_{it}$ (first two columns), the change in the nominal exchange rate Δs_{it} (third and fourth column), the change in the real exchange rate Δe_{it} (fifth and sixth column), and its both components $\Delta\hat{p}_{it}$ and Δs_{it} together (last two columns). The number of stacks refers to the number of bilateral country-pairs included in the panel. For each specification, the table displays the slope coefficient estimates with the corresponding t-statistics (in brackets). Significance at 10%, 5%, and 1% is indicated by *, **, and ***, respectively. The critical values for the t-distribution are 1.645 (10%), 1.960 (5%), and 2.576 (1%).

for this set of countries is available (1986:Q1-2006:Q1).[27] Table 6.10 contains results from fixed-effects panel estimations.[28] Each column in these panels contains results for a different specification.

There are several interesting findings in this table. At least three points are worth to be mentioned. First, the relative output term enters with a significantly positive sign in each specification, suggesting a strong departure from perfect risk-sharing for this group of countries.[29] Second, the negative of the inflation differential enters with a significantly positive sign in all panel estimations. Besides being statistically significant at 1 percent significance

[27] Similar as in the case of the Eurozone countries, all specifications that exclude the nominal exchange rate were estimated for a longer period as well (1970:Q1-2006:Q1). The results differ marginally however, and we do not report them here.

[28] In most specifications for the industrial countries, the Hausman specification test suggests rejection of the null hypothesis of no systematic difference between estimates obtained with fixed-effects and estimates obtained with random-effects. Therefore, only fixed-effects estimations yield consistent estimates in this case.

[29] Comparison of the coefficients in front of the relative output growth term suggest that the departures from perfect risk-sharing are larger for this group of industrial countries relative to the group of Eurozone countries.

level, this effect has an economic meaning as well: it implies values for the relative risk-aversion coefficient in the range 4-9.[30] Third, both nominal and real exchange rate changes enter most panel specifications with a negative sign, though their effect is not statistically significant. In sum, this table gives support to the previous evidence: it suggests that the nominal exchange rate is unrelated to relative consumption growth, while (the negative of) the inflation differential is strongly positively related to the relative consumption growth.

Table 6.11 contains results for the last sample (5): the same group of advanced industrial countries over the Euro-period (1999:Q1-2006:Q1). Table 6.11 presents results from fixed-effects panel estimations and its structure is similar to that of Table 6.10. Several findings in the first panel deserve attention. First, the coefficient in front of the inflation differential is not statistically different from zero in any specification. Second, both nominal and real exchange rate changes have negative signs in all specifications, and their effect is usually (marginally) significant. Third, the coefficient in front of the relative output term is positive and significantly different from zero. Therefore, these findings demonstrate once again that the nominal exchange rate is the main source of Backus-Smith anomaly.

6.4.3 Discussion of the Results

The estimations presented in this section suggested that relative consumption growth is positively related to (the negative of) the inflation component of the real exchange rate. In turn, this finding has important implications about international risk-sharing: it suggests that consumers exploit international relative price changes and increase consumption in situations when their purchasing power is relatively higher (prices they face are relatively lower). This is in line with macroeconomic theory and gives support to the basic condition for efficient risk-sharing in consumption. Moreover, this relation was shown to be surprisingly robust to alternative regression specifications, estimation methods, country samples, and nominal exchange rate regimes. In the case of flexible (floating) nominal exchange rates, the results imply a clear "di-

[30]The figures can be calculated as the reciprocal values of the slope coefficient estimates in front of the inflation differential term ($1/0.23 = 4.34$, $1/0.113 = 8.85$).

Table 6.11: Industrial Countries during the Euro Period 1999:Q1-2006:Q1

Panel Estimates (Fixed Effects)								
$\hat{p}$	0.007	-0.023					-0.003	-0.028
	(0.19)	(-0.67)					(-0.09)	(-0.82)
$\tilde{g}$		0.312		0.313		0.314		0.316
		(8.96)***		(9.04)***		(9.06)***		(9.07)***
s			-0.006	-0.008			-0.006	-0.008
			(-1.46)	(-1.95)*			(-1.45)	(-2.00)**
e					-0.006	-0.009		
					(-1.44)	(-2.03)**		
Obs	840	840	840	840	840	840	840	840
R^2	0	0.09	0	0.09	0	0.09	0	0.09
Stacks	30	30	30	30	30	30	30	30

Note: The table presents results from fixed-effects panel data estimations for 6 industrial countries over the Euro-period 1999:Q1-2006:Q1. The general specification is given by the following regression equation: $\Delta\tilde{c}_{it} = \beta_0 + \beta_1 \Delta x_{it} + u_{it}$, while the specification that allows for partial risk-sharing in consumption is given by: $\Delta\tilde{c}_{it} = \beta_0 + \beta_1 \Delta x_{it} + \beta_2 \Delta\tilde{g}_{it} + u_{it}$. Moreover, each of these specifications is estimated using alternative variables in place of Δx_{it}: the negative of the inflation differential $\Delta\hat{p}_{it}$ (first two columns), the change in the nominal exchange rate Δs_{it} (third and fourth column), the change in the real exchange rate Δe_{it} (fifth and sixth column), and its both components $\Delta\hat{p}_{it}$ and Δs_{it} together (last two columns). The number of stacks refers to the number of bilateral country-pairs included in the panel. For each specification, the table displays the slope coefficient estimates with the corresponding t-statistics (in brackets). Significance at 10%, 5%, and 1% is indicated by *, **, and ***, respectively. The critical values for the t-distribution are 1.645 (10%), 1.960 (5%), and 2.576 (1%).

chotomy" in the behavior of the two components of the real exchange rate. The inflation differential is positively related, while the nominal exchange rate is generally negatively related to relative consumption growth. In turn, this finding suggests that the asset component of the real exchange rate is "disconnected" from the underlying macroeconomic fundamentals (relative consumption), and thereby, generates the consumption-real exchange rate puzzle. Finally, most of these results gain statistical significance when one explicitly accounts for country-pair-specific characteristics (fixed-effects) in the panel regressions.[31]

6.5 Conclusion

This study provides some new empirical evidence about the importance of nominal exchange rate fluctuations for one of the most important anomalies in international macroeconomics: the negative relative consumption growth-real exchange rate correlation (Backus-Smith puzzle). Its main findings can be summarized as follows. First, in accordance with theory and in stark contrast

[31] In general, the specification tests also suggest that this is the only consistent estimation procedure.

to the anomalous behavior first documented by Backus and Smith (1993), the real exchange rate is positively related to relative consumption growth for Eurozone 12 countries after the introduction of the Euro. This is true for the majority of bilateral estimations, but also for the pooled dataset. Second, this finding only applies to the period of fixed nominal exchange rates - the Eurozone countries in the Euro-period, and stays in contrast to all alternative samples with (relatively) flexible nominal exchange rates. Third, when real exchange rate changes for the latter samples are separated into their two components, the regressions results suggest that the Backus-Smith anomaly stems primarily from the nominal exchange rate changes. In fact, there is a clear "dichotomy" between the results for (the negative of) the inflation differential and nominal exchange rate changes. While the first relation is generally positive, and hence, in accordance with theory, the second one is generally negative and implies violation of the basic condition for risk-sharing across countries. Several alternative samples, empirical specifications and procedures support this result. Inevitably, nominal exchange rate behavior appears crucial for understanding the relative consumption-real exchange rate puzzle investigated in this study. Our finding therefore raises an additional puzzle: why is the nominal exchange rate negatively correlated with relative consumption and why does it behave so differently from the inflation differential?

Chapter 7

On the Dichotomy of the Real Exchange Rate

7.1 Introduction

Real exchange rate movements are crucial for the measurement of international risk-sharing (Cole and Obstfeld, 1991; Obstfeld and Rogoff, 1996, 2001; Ravn, 2001; Brandt, Santa-Clara, and Cochrane, 2006). Fluctuations in the real exchange rates imply direct changes in purchasing power across countries, and therefore, determine the direction of payments in an efficient risk-sharing environment. In this way, they create a wedge between marginal utility growth rates. Therefore, the volatility of the real exchange rate relative to the volatility of the underlying fundamentals can be used as a (crude) measure for deviations from (perfect) international risk-sharing (Ravn, 2001; Brandt, Santa-Clara, and Cochrane, 2006).

However, depending on the concrete framework used to quantify the degree of international risk-sharing, real exchange rates appear to be both too volatile (macroeconomic RBC models) and too smooth (asset pricing framework). For example, chapter 6 suggested that real exchange rates are too volatile, and that their anomalous behavior with respect to macroeconomic fundamentals (relative consumption growth) stems primarily from the excessive volatility of their asset component - the nominal exchange rate. On the other hand, chapter 5 demonstrated that under rigid or completely fixed nominal exchange rate regimes, the real exchange rate appears too smooth compared to what is implied by stochastic discount factors retrieved from asset markets. Therefore, the purpose of this chapter is to synthesize some evidence on the importance of the real exchange rate components, starting with the disconnect of real exchange rate volatility.

Backus and Smith (1993) report that bilateral real exchange rates are typically more volatile than what is suggested by the volatility of the relative consumption growth ratio for the corresponding countries. In fact, this anomaly is just one example of the broader class of puzzles concerning the disconnect between exchange rates and their underlying fundamentals that Obstfeld and Rogoff (2001) term *the exchange rate disconnect puzzle.* All phenomena/puzzles in this class point out that (real) exchange rates display excessive volatility and persistence levels compared to what can be justified by the underlying fundamentals in most macroeconomic models: interest rates, output and consumption growth rates, money supply dynamics, etc.[1] In line with these observations, Obstfeld and Rogoff (1996, Chapter 9) suggest that the exchange rate volatility can be understood as one specific manifestation of the excessive volatility present in all asset markets. Therefore, the solution to the exchange rate disconnect puzzle must be sought in a much richer general framework that can give an explanation to the excessive volatility in all these markets (Obstfeld and Rogoff, 1996, Chapter 9).[2]

There are several ways to demonstrate the excessive real exchange rate volatility anomaly using aggregate consumption series.[3] For example, one can choose theoretically "plausible" values for the model parameters, compute moments for some of the model variables, and then compare these moments with those found in actual data. Alternatively, one can directly use actual data for all variables and compare the implied values for the model parameters with

[1] For more extensive discussion about the excessive volatility of the exchange rate relative to fundamentals, see Mark (2001, Chapter 3).

[2] Some theoretical models that produce considerable real exchange rate volatility include Chari, Kehoe, McGrattan (1997, 2002), Thoenissen (2006).

[3] This analysis concentrates only on models that are based on aggregate consumption data. However, the real exchange rate disconnect puzzle can be demonstrated using other theoretical frameworks - for example asset pricing models that use asset returns data only (see Brandt, Santa-Clara, and Cochrane (2006) or Bodenstein (2005)).

the theoretically "plausible" ones. A natural starting point for both types of analysis is the class of real business cycle models. In the presence of perfect international capital markets, these models imply (almost) perfect correlation of consumption growth rates across countries with coefficients of correlation in the vicinity of one. Then, using this theoretically suggested value(s) for the coefficient of correlation between consumption growth rates, these models can reconcile the actual real exchange rate volatility with the relative consumption growth volatility only if the coefficient of relative risk-aversion is extremely high, typically in the order of 100. As the implied value for the risk-aversion coefficient is much higher than what is considered "plausible"(typically not more than 10), this finding gives rise to (one form of) the volatility disconnect puzzle.[4]

However, a large body of literature documents that correlations of consumption growth rates across (OECD) countries are much lower than unity and even lower than correlations of output growth rates (see, for example, Backus, Kehoe, and Kydland, 1992). In turn, these studies typically find values of about 0.3 for the coefficient of correlation for consumption growth rates of the most advanced, OECD countries (Pakko, 2004). Using these empirically calculated consumption correlations, the implied coefficient of risk-aversion is considerably lower and usually within the upper bound of what is considered reasonable (in the order of 5-10).[5] However, this raises the issue of the consumption correlation puzzle, which appears closely related to the excessive volatility of the real exchange rate compared to relative consumption growth. Actually, possible solutions to the former puzzle (explaining why consumption growth is much less correlated across countries than theoretically expected) might help in resolving the latter puzzle as well (by bridging the large gap

[4]The economic literature is not unanimous and suggests different ranges about the theoretically "plausible" values for the coefficient of relative risk-aversion. For example, several seminal studies place it in the range between one and two: Arrow (1971), Friend and Blume (1975), Kydland and Prescott (1982). On the other hand, Kandel and Stambaugh (1991) argue that even values of about 30 seem reasonable. This study follows a "middle way" in accordance with Mehra and Prescott (1985, 1988) and Mankiw and Zeldes (1991) and views values of 10 as the upper-bound of what can be considered "plausible".

[5]For calculations using empirically found values for the consumption correlation coefficient see Bodenstein (2005), for example.

between theoretically plausible and empirically calculated values for the risk coefficient). Note though that the use of empirical consumption correlations attenuates the excessive volatility puzzle, but does not completely solve it.

A large body of empirical literature suggests that the major part of real exchange rate volatility can be attributed to nominal exchange rate fluctuations (Mussa, 1986; Baxter and Stockman, 1989; Flood and Rose, 1999; Obstfeld and Rogoff, 1996, 2001). Therefore, one way to find an explanation of the exchange rate disconnect puzzle is by looking at the contribution of each of its components: the nominal exchange rate fluctuations, and the inflation differential changes. In turn, the establishment of monetary union and the adoption of common currency by 11 European countries in 1999 (12 since June 2000, and 13 since January 2007) provides a “natural” test about the importance of the nominal exchange rate for the volatility disconnect puzzle.

Using data for the Eurozone countries, the first part of this study tries to give an answer to the following question: Does the elimination of the nominal exchange rate solve the exchange rate disconnect puzzle? The main findings suggest that the elimination of the nominal exchange rate seems to “over-fix” the exchange rate disconnect puzzle as the implied value for the coefficient of relative risk-aversion turns “implausibly” low, even lower than what is implied by logarithmic utility function, for example.

In the second part of the chapter, we relate these findings for the real exchange rate volatility-disconnect puzzle with the broader “disconnect” within the real exchange rate suggested in the previous chapters (chapters 5 and 6). Focusing on the UK-USA country-pair, we show that a clear “dichotomy” exists between the macroeconomic and the asset side of the real exchange rate.

The rest of this chapter is organized as follows. The first part of the chapter (section 7.2) deals with the volatility disconnect puzzle. Section 7.2.1 presents the theoretical framework and derives empirically testable versions of the main conditions. Section 7.2.2 briefly describes the dataset and the empirical strategy followed, while section 7.2.3 presents the main results. The second part (section 7.3) concentrates on the dichotomy between the two sides/components of the real exchange rate. Section 7.3.1 describes the dataset and empirical strategy, while sections 7.3.2 and 7.3.3 present the main findings. Finally, section 7.4 gives some concluding remarks.

7.2 Volatility Disconnect Puzzle

7.2.1 Theoretical Framework

This section starts with a brief theoretical note on the puzzle investigated in the first part of this chapter: the excessive volatility of the real exchange rate compared to macroeconomic fundamentals. As in the case of the Backus-Smith puzzle (elaborated on in chapter 6), we start the discussion with the same condition for efficient risk-sharing across countries in the presence of real exchange rate fluctuations:

$$\frac{e_{t+1}}{e_t}\frac{u'(c_{t+1})}{u'(c_t)} = \frac{u'(c^*_{t+1})}{u'(c^*_t)} \tag{7.1}$$

which, after taking logarithms from both sides and a slight re-arrangement, results in the following equation:

$$\ln\frac{e_{t+1}}{e_t} = \ln\frac{u'(c^*_{t+1})}{u'(c^*_t)} - \ln\frac{u'(c_{t+1})}{u'(c_t)} \tag{7.2}$$

The utility function (and hence the change in the intertemporal marginal rate of substitution) in this equation is given in general form. In order to implement this equation empirically, a specific form of utility function has to be assigned. We assume that preferences and identical across countries and can be described by the constant relative risk aversion (CRRA) utility function (following the example with the Backus-Smith condition):

$$u(c_t) = \frac{(c_t)^{1-\gamma}}{1-\gamma} \tag{7.3}$$

where γ stands for the coefficient of relative risk aversion. Using this CRRA form for the utility function, condition (7.2) can be rewritten as follows:

$$\ln\left(\frac{e_{t+1}}{e_t}\right) = \gamma\ln\left(\frac{c_{t+1}}{c_t}\right) - \gamma\ln\left(\frac{c^*_{t+1}}{c^*_t}\right) \tag{7.4}$$

Taking variances of both sides of equation (7.4), we obtain the following expression:

$$\sigma^2\Big(\ln\frac{e_{t+1}}{e_t}\Big) = \gamma^2\sigma^2\Big(\ln\frac{c_{t+1}}{c_t}\Big) + \gamma^2\sigma^2\Big(\ln\frac{c^*_{t+1}}{c^*_t}\Big) - \\ - 2\gamma^2\sigma\Big(\ln\frac{c_{t+1}}{c_t}\Big)\sigma\Big(\ln\frac{c^*_{t+1}}{c^*_t}\Big)\rho\Big(\ln\frac{c_{t+1}}{c_t}, \ln\frac{c^*_{t+1}}{c^*_t}\Big) \tag{7.5}$$

where $\rho\Big(\ln\frac{c_{t+1}}{c_t}, \ln\frac{c^*_{t+1}}{c^*_t}\Big)$ is the coefficient of consumption growth correlation between the two countries. Therefore, equation (7.5) relates the variance of the bilateral real exchange rate with the variances of consumption growth in the corresponding countries, the coefficient of consumption growth correlation, and the coefficient of relative risk aversion. In turn, the coefficient of relative risk aversion can be expressed in terms of the other variables as follows:[6]

$$\gamma = \frac{\sigma\Big(\ln\frac{e_{t+1}}{e_t}\Big)}{\sqrt{\sigma^2\Big(\ln\frac{c_{t+1}}{c_t}\Big) + \sigma^2\Big(\ln\frac{c^*_{t+1}}{c^*_t}\Big) - 2\sigma\Big(\ln\frac{c_{t+1}}{c_t}\Big)\sigma\Big(\ln\frac{c^*_{t+1}}{c^*_t}\Big)\rho\Big(\ln\frac{c_{t+1}}{c_t}, \ln\frac{c^*_{t+1}}{c^*_t}\Big)}} \tag{7.6}$$

or expressed in terms of rates of change in discrete time:

$$\gamma = \frac{\sigma^2(\Delta e)}{\sqrt{\sigma^2(\Delta c) + \sigma^2(\Delta c^*) - 2\sigma(\Delta c)\sigma(\Delta c^*)\rho(\Delta c, \Delta c^*)}} \tag{7.7}$$

Most empirical studies demonstrate that only if this coefficient is very high (typically about 100, Obstfeld and Rogoff, 1996; 2001), theoretical macroeconomic models (with $\rho = 1$) match the data. In fact, finding implausibly high implied values for the coefficient of relative risk-aversion in this theoretical framework is equivalent to measuring excessively high volatility of the real exchange rate compared to economic fundamentals. This is not surprising since complete markets models imply perfect correlation between the underlying consumption streams, and therefore, very smooth real exchange rate.

There are two basic steps to bridge this gap between implications from the theoretical models and the empirical evidence. First, the use of empirically found consumption correlations "relaxes" the strict bounds on real exchange

[6] In this framework, we do not assume that consumption growth rates are equally volatile across countries, but rather present the general case and then calculate them from actual data. For the corresponding equations under the simplifying assumption that consumption growth rates are equally volatile across countries see Bodenstein (2005), for example.

rate volatility imposed in the canonical macroeconomic models. Second, the gap might be bridged through a dramatic (exogenous) fall in real exchange rate volatility. Therefore, if the real exchange rate volatility puzzle indeed disappears with the elimination of the nominal exchange rate, then the implied values for the coefficient of relative risk-aversion should be in the "plausible" range (usually in the range $1 - 10$).

7.2.2 Data and Empirical Strategy

Data Sources

The analysis is based on the same dataset used in chapter 6. It includes quarterly observations for all 12 Eurozone countries as well as six major industrial countries outside the Eurozone: Australia, Japan, Sweden, Switzerland, UK, and USA over the period $1986 : Q1 - 2006 : Q1$. The following four macroeconomic time series are included in the dataset: real per capita consumption, real per capita gross domestic product, nominal exchange rates and inflation rates. Data on consumption and gross domestic product comes from the OECD Main Economic Indicators database. Data on nominal exchange rates comes from GTIS database. For all countries in the dataset, we calculate quarterly changes in the nominal exchange rates (calculated through the US/domestic currency GTIS exchange rate series). The inflation rate series are calculated as quarterly changes in the general CPIs (alternatively NADJ). All data is retrieved from Datastream.

Data Samples

This study is primarily interested in empirically investigating whether the volatility-disconnect puzzle decreases/disappears for the 12 countries of the Eurozone during the Euro period. Nonetheless, in order to capture the importance of the elimination of the nominal exchange rates within the Eurozone, we compare it with four other samples (groups of countries/time periods). All of these samples are shown in Table 7.1. Sample (1) refers to the sample of primary interest, i.e. the countries of the Eurozone during the Euro period (1999-2006). First, the results for this sample are compared with the results for the same group of countries in different time periods. Thus, sample (2)

Table 7.1: Samples and Time Periods

Sample	Countries	Time period
(1)	Eurozone	1999:Q1-2006:Q1
(2)	Eurozone countries	1986:Q1-1998:Q4
(3)	Eurozone countries	1995:Q1-1998:Q4
(4)	Industrial countries	1986:Q1-2006:Q1
(5)	Industrial countries	1999:Q1-2006:Q1

covers a long period (1986-1998) and sample (3) covers a shorter period before the introduction of the Euro (1995-1998). This specific division of the entire time period was dictated by data availability issues. The former sample (2) covers the longest period for which data was available, but this panel dataset is strongly unbalanced. The latter sample (3) covers a much shorter period with a strongly balanced panel dataset. Second, the results from sample (1) are compared with two samples for a group of (major) industrial countries: sample (4) covering the entire time period (1986-2006), and sample (5) covering the Euro period only (1999-2006).

Summary Statistics

Since the empirical analysis about the volatility-disconnect puzzle is based on the same dataset that is used in chapter 6, we do not repeat or discuss any summary statistics for these data series here. Instead, they can be found in the tables in section 6.3.3.

7.2.3 Results

This section presents the main findings from the empirical analysis about the volatility-disconnect puzzle. The presentation of these findings comprises two parts. First, we present some visual evidence about the importance of the (elimination of the) nominal exchange rate for the relative consumption-real exchange rate volatility comparison. Second, we present calculations about the implied values of the coefficient of relative risk-aversion γ for each of the five samples.

Volatility Comparison: Visual Evidence

This section starts with some visual evidence on the real exchange rate volatility puzzle. For this purpose, Figure 7.1 shows several comparisons between relative consumption growth and real exchange rate volatility (both measured as standard deviations). Each of the six scatterplots presented in this figure refers to a separate data sample:

- (upper-left) All 12 Eurozone countries (1999:Q1-2006:Q1, 66 observations);
- (upper-right) Eurozone countries excluding Greece and Luxembourg (1999:Q1-2006:Q1, 45 observations);
- (middle-left) All 12 Eurozone countries (1986:Q1-1998:Q4, 66 observations);
- (middle-right) All 12 Eurozone countries (1995:Q1-1998:Q4, 66 observations);
- (lower-left) Industrial countries (1986:Q1-2006:Q1, 15 observations)
- (lower-right) Industrial countries (1999:Q1-2006:Q1, 15 observations)

The total number of observations per scatterplot is given by the number of (bilateral) country-pair combinations in the corresponding sample. For a sample/set of size N the number of bilateral combinations is $N(N-1)/2$.[7] The figure shows an interesting pattern: the upper plots for the Eurozone countries after the introduction of the Euro indicate that the real exchange rate was equally volatile as the consumption growth rate. In this period, the nominal exchange rate is eliminated, and therefore, the real exchange rate volatility equals the volatility of the inflation differential. Hence, once the nominal exchange rate does not play a role, the (excessive) real exchange rate volatility anomaly disappears: the real exchange rate between two countries is

[7] For the samples including Eurozone 12 it is $12(12-1)/2 = 66$, for the Eurozone sample without Greece and Luxembourg it is $10(10-1)/2 = 45$, and for the sample of 6 Industrial Countries it is $6(6-1)/2 = 15$.

comparable to the volatility of the underlying fundamentals, i.e. the relative consumption growth.[8]

In order to demonstrate that this puzzle only disappears when the nominal exchange rate is (irrevocably) fixed, we contrast the results for the Eurozone countries after the introduction of the Euro with four alternative samples. The scatterplots in the middle and lower rows of Figure 7.1 show evidence about these four alternative samples. The two plots in the middle row of the figure refer to the Eurozone countries before the introduction of the Euro. In fact, almost all data points in these plots lie (high) above the 45 degrees line, suggesting that the real exchange rate was considerably more volatile compared to the relative consumption growth rate. Furthermore, the plots in the last row in this figure refer to the set of industrial countries (samples (4) and (5)). The disconnect between real exchange rate and relative consumption growth volatility, is even stronger than before. In fact, most data points lie very close to the ordinate-axis, and high above the 45 degrees line. Conclusively, these plots very clearly demonstrate the importance of the volatility-disconnect puzzle: the real exchange rate is much more volatile and disconnected from economic fundamentals (Obstfeld and Rogoff, 2001).

Implied Values for γ

The previous figure presented some visual evidence about the importance of (the elimination of) the nominal exchange rate for the (in)existence of the real exchange rate volatility puzzle. Now, we go one step further and present some analytical results. For this purpose, we calculate values for the implied coefficient of relative risk aversion according to equation (7.8):

$$\gamma = \frac{\sigma\left(\ln \frac{e_{t+1}}{e_t}\right)}{\sqrt{\sigma^2\left(\ln \frac{c_{t+1}}{c_t}\right) + \sigma^2\left(\ln \frac{c^*_{t+1}}{c^*_t}\right) - 2\sigma\left(\ln \frac{c_{t+1}}{c_t}\right)\sigma\left(\ln \frac{c^*_{t+1}}{c^*_t}\right)\rho\left(\ln \frac{c_{t+1}}{c_t}, \ln \frac{c^*_{t+1}}{c^*_t}\right)}} \tag{7.8}$$

where $\sigma(\ln \frac{e_{t+1}}{e_t})$ stands for (overall) real exchange rate standard deviation (in samples 2-5) and $\rho(\ln \frac{c_{t+1}}{c_t}, \ln \frac{c^*_{t+1}}{c^*_t})$ stands for the empirically-found

[8]The first plot (upper-left) contains several outliers, i.e. extremely high observations for the relative consumption growth volatility that refer to Greece and/or Luxembourg after the introduction of the Euro. The second plot excludes these extreme observations.

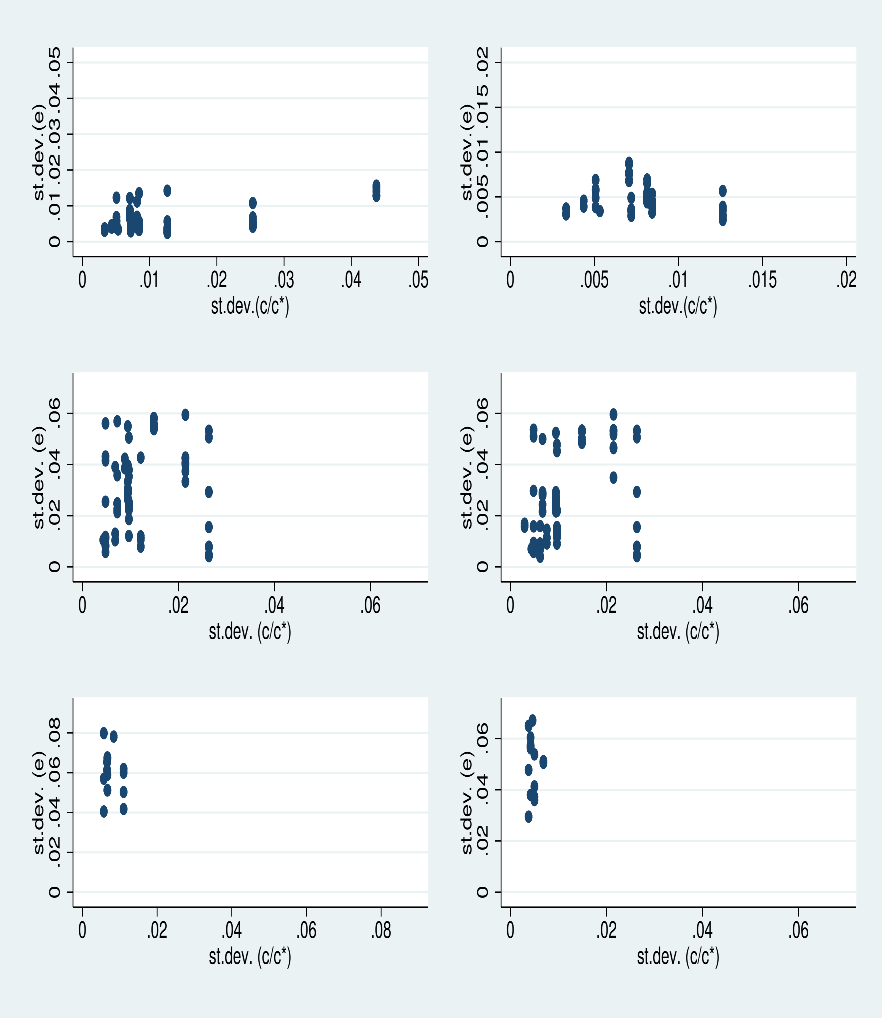


Figure 7.1: Volatility Comparison: Relative Consumption Growth v Real Exchange Rate

Note: The figure presents a comparison between real exchange rate volatility σ_e (on the vertical axis) and relative consumption growth volatility $\sigma_{c/c*}$ (on the horizontal axis). Volatility is measured as quarterly standard deviation. Each scatterplot refers to a different sample: Eurozone (1999:Q1-2006:Q1, upper-left); Eurozone exc.Greece and Luxembourg (1999:Q1-2006:Q1, upper-right); Eurozone (1986:Q1-1998:Q4, middle-left); Eurozone (1995:Q1-1998:Q4, middle-right); Industrial (1986:Q1-2006:Q1, lower-left), Industrial (1999:Q1-2006:Q1, lower-right).

Table 7.2: Summary Statistics: Implied Values for the Relative Risk-Aversion Coefficient (γ)

Group	Obs	Mean	Std. Dev.	Min	Max
Eurozone (1)	66	0.482	0.230	0.133	1.116
Eurozone (2)	66	2.267	1.392	0.127	6.821
Eurozone (3)	66	2.124	1.456	0.143	6.248
Industrial Countries (4)	15	5.908	1.625	3.863	9.179
Industrial Countries (5)	15	8.221	2.506	4.292	11.959

Note: The table contains summary statistics for the implied values of the coefficient of relative risk-aversion γ calculated according to equation (7.8). Each row refers contains figures for a separate sample from Table 7.1. All implied values are calculated using quarterly observations for the corresponding period.

coefficient of correlation between the consumption series. Since the nominal exchange rate is fixed in sample (1), $\sigma(\ln \frac{e_{t+1}}{e_t})$ refers to the standard deviation of the inflation differential. Table 7.2 presents results from the calculations. Each row in this table refers to a separate data sample ((1)-(5)). The second column gives the total number of observations used in the calculations.[9] The other columns present the results of the calculations: the mean value for the relative risk-aversion coefficient, its standard deviation, and its two extremes (minimum and maximum). In general, these figures indicate high variability in the calculation of implied values for the risk-aversion coefficient.

The third column in this table, which contains the average implied values for the coefficient of relative risk-aversion, conveys the main messages from this exercise. First, the implied values for the coefficient of relative risk-aversion are much lower for the Eurozone compared to those for the group of industrial countries. Second, the implied value for the Eurozone 12 countries after the introduction of the Euro (sample (1)) is very low. In fact, the average value for this sample (0.4822) is even lower than unity, the value for the coefficient of relative risk-aversion that refers to logarithmic CRRA utility function.

The remaining three columns just strengthen these conclusions. The fig-

[9]Again, the total number of bilateral combinations for a sample/set of size N is calculated by $N(N-1)/2$. For the Eurozone samples ((1), (2), and (3)), we base the calculations on all possible 66 ($12(12-1)/2 = 66$) bilateral combinations, while for the set of 6 advanced industrial countries (samples (4) and (5)) on all possible 15 ($6(6-1)/2 = 15$) observations/combinations.

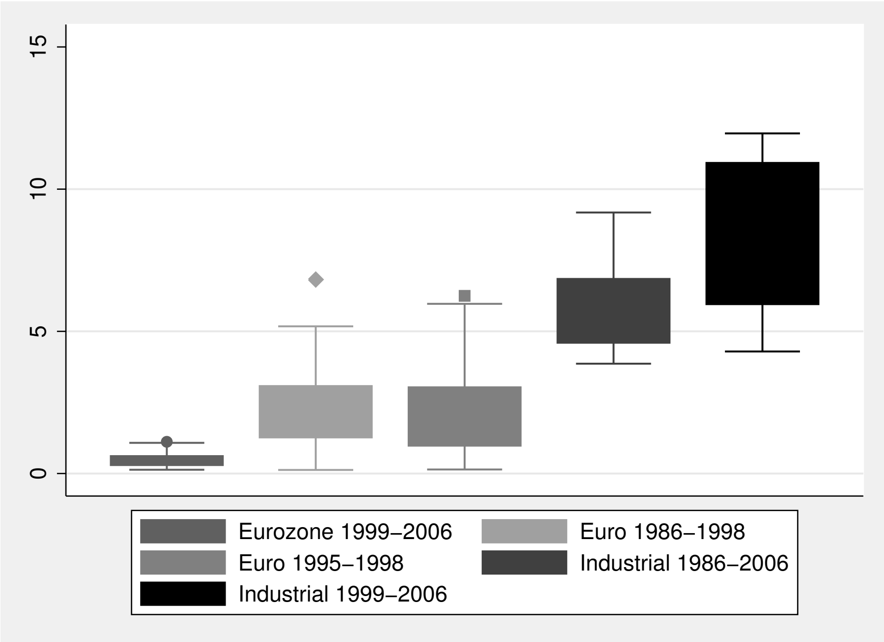


Figure 7.2: Implied Values for the Coefficient of Relative Risk-Aversion (γ)

Note: This figure presents box plots for the implied values of the coefficient of relative risk-aversion (γ) for each of the 5 samples. The upper hinge refers to the 75th percentile, while the lower hinge refers to the 25th percentile in each box plot. Hence, the box contains all values that fall in the interquartile range. The whiskers contain values that fall within two standard deviations from the mean, while all values outside this range are marked as outliers (outside values).

ures for the standard deviation (fourth column) suggest that the implied values (estimates) for the coefficient of relative risk-aversion in the Eurozone samples are much less dispersed compared to those for the samples of industrial countries. Furthermore, the extreme values in the last two columns offer additional evidence about the systematic difference between the implied coefficients for different samples. In fact, even the lowest value for a pair of industrial countries lies above the average value for Eurozone country pairs.[10]

Figure 7.2 gives a graphical representation of the results from Table 7.2. It presents several boxplots, each containing implied values for the coefficient of relative risk-aversion for one of the five samples. The visual evidence in

[10]Moreover, the maximum implied value for a country pair in sample (1) is very close to unity (1.116), suggesting logarithmic utility.

this figure just strengthens the conclusions from Table 7.2: the implied values for the industrial countries (samples (4) and (5)) are higher compared to the Eurozone before the introduction of the Euro (samples (2) and (3)), which in turn, are considerably higher compared to the Eurozone after the introduction of the Euro (sample (1)). In fact, most of the implied values for the industrial countries correspond very closely to those found in the empirical literature using empirical consumption correlation figures, falling in the range 5-10. Furthermore, the entire interquartile ranges for these countries is above or very close to the maximum values (outliers) for the Eurozone countries (samples (2) and (3)). Finally, a similar conclusion applies to the group of Eurozone countries after the introduction of the common currency: in fact, all values for this sample fall below the lower hinge (25th percentile) for the same group of countries before the introduction of the Euro.

Discussion about the Results

The results presented above suggest that the volatility-disconnect puzzle can be bridged from two sides: first, by using empirically-found consumption correlations that are much lower than theoretically-suggested ones; second, by eliminating the nominal exchange rate - the main source of real exchange rate volatility. The first step - use of empirically-found consumption correlations - brought the results for the coefficient of relative risk-aversion within the upper-bound of what is usually considered “plausible”. Hence, this finding demonstrates that the severity of the volatility-disconnect puzzle essentially results mainly from the use of theoretically-suggested rather than empirically-found consumption correlations. In this way, the volatility-disconnect might be very closely related to the consumption-correlation puzzle - explanations for the latter puzzle will necessarily help in resolving the former puzzle as well. The second step - elimination of the nominal exchange rate - seems to fix or even “over-fix” the volatility-disconnect puzzle. First, the reasonable values for the coefficient of relative risk-aversion found for the Eurozone countries before the introduction of the Euro (sample 2 and 3) broadly reflect the stabilization (control) of the nominal exchange rates among these countries in the pre-Euro period. In this way, the (quasi) elimination of the asset component of the real exchange rate largely fixes the volatility-disconnect puzzle.[11] Second, the complete elimination of the nominal exchange rate after the introduction of the Euro even seems to over-fix the puzzle. This is evident for sample (1) (Eurozone countries in the Euro-period), where it leads to very low values for the coefficient of relative risk-aversion.[12]

In sum, the implied values for the coefficient of relative risk-aversion suggested that both the use of empirical consumption correlations and the (quasi) elimination of the nominal exchange rate seem to contribute to the resolution of the volatility-disconnect puzzle.[13] Moreover, only the combination of both factors together: empirical correlations with complete elimination of the nominal exchange rate tends to over-fix the puzzle. In this case, the specific combination of factors might even suggest a possible anomaly in the opposite direction: real exchange rate appears too smooth relative to fundamentals (sample 1).

Implications about Price Flexibility in the Eurozone

In the previous section, the elimination of the nominal exchange rate was seen to cause a sharp and permanent fall in the flexibility of intra-Eurozone real exchange rates. With respect to the calculations in the previous section, this point can be illustrated with a comparison of the implied values for the coefficient of relative risk-aversion before and after the introduction of the Euro. Figure 7.2 and Table 7.2 showed that the average implied value dropped from slightly above 2 in the former period to about 0.5 in the latter period. In turn, this implies that intra-Eurozone real exchange rates became considerably smoother after the introduction of the common currency.[14]

[11]The implies values in samples (2) and (3) are about 2 on average, which is considered reasonable for most studies mentioned in the introduction.

[12]Most studies suggest that values below unity (logarithmic function) are probably too low to be considered reasonable (Kocherlakota, 1996; Mehra and Prescott, 1985, 1988; Mankiw and Zeldes, 1991).

[13]The last option to be investigated is the use of theoretically-suggested consumption correlations with eliminated nominal exchange rate. In that case, the average implied value for the coefficient of relative risk-aversion for the Eurozone after the introduction of the Euro is about 3. The complete results for all possible combinations are available upon request.

[14]If the denominator in equation (7.8) did not experience a dramatic change, then the considerable fall in the implied values for γ must come from the decrease in the numerator.

This observation can be related to the evolution of price flexibility in the Eurozone. If the introduction of the Euro does not (instantly) lead to significant, structural changes in the real economy, then no abrupt changes in intra-Eurozone real exchange rate flexibility should be expected. In this case, domestic, intra-Eurozone prices should get more variable (flexible), compensating for the loss of the nominal exchange rate as an adjustment channel. In the context of the framework presented here, domestic prices should have "taken over" the adjustment role of the nominal exchange rate, thereby assuring about the same level of real exchange rate flexibility, and therefore, similar implied values for the coefficient of relative risk-aversion. However, the calculations in this chapter show that this is not the case in the Eurozone. In fact, there is a dramatic fall in the implied values for the coefficient of relative risk-aversion suggesting that domestic prices in the Eurozone are far from flexible enough to compensate for the loss of the nominal exchange rate as an adjustment channel after the introduction of the Euro.

Furthermore, this finding can be related to the main question in the analysis: the importance of the nominal exchange rate rate for the resolution of the volatility-disconnect puzzle. The calculations for the pre-Euro period suggest that the volatility-disconnect puzzle disappears when nominal exchange rates are stabilized, but not entirely eliminated - reasonable values for the coefficient of relative risk-aversion can match real exchange rate with relative consumption volatility (about 2 in samples 2 and 3).

However, the complete elimination of the nominal exchange rates after the introduction of the Euro was not followed by a proportionate increase in relative price variability/flexibility in order to keep the impled values for γ at similar, reasonable levels. Instead, the significant drop in these implied values (below the value that corresponds with logarithmic utility) suggests a possible puzzle in the opposite direction: once deprived of its asset component, the real exchange rate appears too smooth (not too volatile) compared to macroeconomic fundamentals (relative consumption growth in this case).

7.3 Macroeconomic and Asset Sides of the Real Exchange Rate

The findings in the previous section demonstrated that, depending on the underlying nominal exchange rate regime, the volatility of the real exchange rate can be higher or lower compared to what is implied by fundamentals. On one hand, the real exchange rate seems to be very volatile when nominal exchange rates are (relatively) flexible (samples 4 and 5). On the other hand, its volatility seems to be quite low when nominal exchange rates are completely fixed as in the case of the Eurozone (sample 1).

These observations suggest that the behavior of the real exchange rate crucially depends on the relative importance of its two components: the nominal exchange rate and the inflation differential. As long as it is allowed to float (relatively) freely, the nominal exchange rate clearly dominates the inflation differential and determines the behavior of the real exchange rate.

The first part of this chapter demonstrated that the volatility-disconnect applies mainly to floating exchange rate regimes. Moreover, chapter 6 demonstrated that floating nominal exchange rates are the main source of another persistent puzzle in international macroeconomics: the consumption-real exchange rate correlation (Backus-Smith) puzzle. In both of these cases, the nominal exchange was viewed as too volatile and therefore disconnected from the underlying fundamentals. On the other hand, chapter 5 showed that under rigid nominal exchange rate regimes, the real exchange rate appears too smooth compared to what stochastic discount factors retrieved from asset markets imply. In sum, all of these findings point towards a broad disconnect between the asset and macroeconomic sides of the real exchange.

Therefore, in this section, we further investigate this "dichotomy" in the real exchange rate. Focusing on the UK-USA country-pair over the period $1975 : Q1 - 2006 : Q1$, we provide evidence on several relationships of (the components of) the real exchange rate with the macroeconomic fundamentals and asset markets returns.

7.3.1 Data Issues and Empirical Strategy

We focus on the country-pair UK-USA because of several reasons. First, this country-pair has (probably) the most accurate statistics on all data-series included in the analysis over a very long time period. Second, these two countries are included in most empirical studies that investigate the puzzles related to the real exchange rate. Finally, the empirical analyses in chapters 4, 5, and 6 and the first part of this chapter all included these two countries.

The dataset consists of quarterly observations on macroeconomic and financial time-series already used in the previous chapters: nominal exchange rates, inflation rates, real per capita consumption growth rates, and stock market returns. Nominal exchange rates series come from the GTIS database, while inflation rates are calculated as rates of change in the general CPIs. Data on real per capita consumption comes from the OECD Main Economic Indicators database. All these macroeconomic series are also used in the empirical analysis in chapter 6 and in the first part of this chapter. Asset returns are calculated as the quarterly rates of change in the leading stock market indices for the corresponding countries: S&P 500 for the USA, and FTSE ALL (or FTSE 100) for the UK.[15] The same series for asset market returns were also used in the empirical analysis in chapters 4 and 5. All data-series are retrieved from Datastream.

The empirical analysis proceeds as follows. First, we present a comparison between values for the coefficient of relative risk-aversion implied by consumption growth and values implied by asset markets returns. Second, we present simple correlations between the components of the real exchange rate and the underlying macroeconomic fundamentals or asset returns. Third, we formally test for several "parity" relations involving these variables.

[15]The calculations for the complete period ($1975:Q1-2006:Q1$) use FTSE ALL, because the series for FTSE 100 start only in $1978:Q2$. The results from the calculations and the regression estimations are very robust to the choice of particular stock market index (FTSE ALL or FTSE 100).

7.3.2 Implied Values for γ

Table 7.3 presents a comparison between values for the coefficient of relative risk-aversion γ implied by macroeconomic data (consumption growth) and values implied by financial markets data (asset markets returns). All values are calculated following equation (7.8) over the period $1975 : Q1 - 2006 : Q1$.[16] The left half of the table contains results from calculations that use consumption growth rates as underlying fundamentals, while the right half of the table contains results from calculations based on asset markets returns. For each of these two approaches, we calculate implied values for γ when either volatility of the (complete) real exchange rate, or the volatility of one of its components is used as numerator in equation (7.8). The column marked with e employs (total) real exchange rate volatility, while the columns marked with s and $\hat{p}$ employ nominal exchange rate and inflation differential volatility, respectively.[17] For each of these specifications, we calculate the required values for γ for alternative correlations between the underlying processes (consumption growth or stock markets returns). The first row refers to the case of perfect correlation ($\rho = 1$), while the second row uses empirically found values for the coefficient of correlation ρ.

Several findings in Table 7.3 deserve attention. First, when the underlying processes are assumed to be perfectly correlated ($\rho = 1$), then equation (7.8) implies very high values for γ. If consumption growth in one country (UK) is very strongly (perfectly) correlated with consumption growth in the other country (USA), then their real exchange rate is not expected to move a lot, unless consumers in these countries are extremely risk-averse.[18] Therefore, the implied values for $\rho = 1$ are always higher than in the case of empirically found ρ. Second, the real and the nominal exchange rates are much more volatile than consumption growth - a fact reflected in (relatively) higher values for γ in the left-hand panel. Third, the results for the inflation differential $\hat{p}$ in the

[16] Calculations over shorter periods, e.g. over the Euro-period $1999 : Q1 - 2006 : Q1$ give very similar results. These results are available upon request.

[17] All volatility figures are calculated as quarterly standard deviations for the corresponding series.

[18] This is a standard result from the Euler's condition with constant relative risk aversion (CRRA) utility functions. See equation (7.4) for example.

Table 7.3: Implied Values for the Coefficient of Risk Aversion

Consumption Growth				Asset Markets Returns			
ρ	e	s	$\hat{p}$	ρ	e	s	$\hat{p}$
1975:Q1-2006:Q1							
1.000	11.328	11.198	2.753	1.000	2.552	2.523	0.620
0.125	4.550	4.498	1.106	0.696	0.733	0.725	0.178

Note: The table presents figures for the implied value for the coefficient of relative risk-aversion γ calculated according to the following equation: $\gamma = \frac{\sigma(\ln \frac{a_{t+1}}{a_t})}{\sqrt{\sigma^2(\ln \frac{b_{t+1}}{b_t})+\sigma^2(\ln \frac{b^*_{t+1}}{b^*_t})-2\sigma(\ln \frac{b_{t+1}}{b_t})\sigma(\ln \frac{b^*_{t+1}}{b^*_t})\rho(\ln \frac{b_{t+1}}{b_t},\ln \frac{b^*_{t+1}}{b^*_t})}}$ Depending on the exact specification used, a and b in this equation refer to different variables. The left panel contains implied values calculated using consumption growth rates. For this setting, a in the above equation refers to the real exchange rate e or one of its components $\hat{p}$ and s, while b in the above equation refers to real per capital consumption c. The right panel contains implied values calculated using assets markets returns. For this setting, a in the above equation refers to the real exchange rate e or one of its components $\hat{p}$ and s, while b in the above equation refers to the level of the stock market index R in the corresponding country. All figures are calculated using quarterly observations for the corresponding variables over the period $1975 : Q1 - 2006 : Q1$. Results from two calculations are presented for each specification: the first assumes 'perfect' correlation between the underlying processes ($\rho = 1$), while the second uses the empirical correlation.

left-hand panel (macroeconomic fundamentals) are surprisingly similar to the figures for the (real or nominal) exchange rate in the right-hand panel (asset markets returns). Finally, it is worth noting that the results are very similar whether the real or the nominal exchange rate is compared to consumption growth or asset markets returns. This finding is not surprising as it reflects the fact that nominal exchange rate constitutes by far the largest (dominant) component of real exchange rate volatility.

The evidence presented in this table serves two purposes. First, it shows how large differences exist between values implied by macroeconomic data and values implied by financial markets data. Second, it demonstrates how similar these two approaches are when the macroeconomic component of the real exchange rate is compared to the underlying macroeconomic fundamentals, while the asset component of the real exchange rate is compared to the underlying asset markets returns. For example, when inflation differential volatility is compared to the volatility of the underlying consumption processes, the implied value for γ is 2.753 (perfect consumption correlation assumed) or 1.1 (empirically-found correlation used).[19] These values are surprisingly similar

[19] Both values seem "plausible" w.r.t. the discussion in the first part of this chapter.

to those obtained in the right-hand panel - when (nominal or real) exchange rate volatility is compared to asset markets returns. The corresponding figures are 2.552 (2.523) (perfect returns correlation assumed) and 0.733 (0.725) (empirically-found correlation used). Although large differences exist between them, both sides imply almost identical values for the coefficient of relative risk-aversion. In turn, these results suggest that the behavior of the macroeconomic part of the real exchange rate relative to its macroeconomic fundamentals is extremely similar with the behavior of the asset part of the real exchange rate relative to its underlying asset markets returns.

7.3.3 Testing for Parity Conditions

The previous section suggested that a "disconnect" exists between the two components of the real exchange. In this section, we further analyze the importance of this dichotomy for several parity conditions involving the (real) exchange rate. For this purpose, we concentrate on five variables for the UK-USA country pair: real and nominal exchange rate changes, inflation differential, stock market returns differential, and relative consumption growth rate. The nominal exchange rate s is defined as the number of pounds sterling per US dollar (£/\$). Therefore, Δs refers to the change in the nominal exchange rate - positive values indicate appreciation of the US dollar. Relative consumption growth is defined as the difference between consumption growth in UK and consumption growth in the USA: $\Delta\tilde{c}_{t+1} \equiv \Delta(c_{t+1}^{UK}) - \Delta(c_{t+1}^{USA})$. In a similar way, we define the stock market returns differential: $\Delta\tilde{r}_{t+1} \equiv \Delta(r_{t+1}^{UK}) - \Delta(r_{t+1}^{USA})$ and the inverted inflation differential: $\Delta\hat{p}_{t+1} \equiv \Delta(P_{t+1}^{USA}) - \Delta(P_{t+1}^{UK})$.[20] Finally, using this notation, the real exchange rate is defined as follows: $e = s\frac{P^{USA}}{P^{UK}}$, and its change $\Delta e = \Delta s + \Delta\hat{p}$.

The correlations among these variables are presented in Table 7.4. If theoretical parity conditions hold, than all correlations in this table should be positive and close to unity.[21] However, the results in this table suggest that only half of the empirical correlations have the "correct" positive sign. In fact,

[20] Similar as in chapter 6, $\Delta\hat{p}_{t+1}$ is defined as the negative of the inflation differential, so that it is positively related to the change in the real exchange rate.

[21] All variables included in this correlation matrix are defined in such a way that parity conditions imply positive correlations among them.

Table 7.4: Correlation Matrix

	Δs	$\Delta\tilde{r}$	$\Delta\tilde{c}$	$\Delta\hat{p}$	Δe
Δs	1				
$\Delta\tilde{r}$	0.199	1			
$\Delta\tilde{c}$	-0.242	-0.114	1		
$\Delta\hat{p}$	-0.089	-0.047	0.277	1	
Δe	0.972	0.187	-0.175	0.148	1

Note: The table presents figures for the coefficient of (bilateral) correlation between the change in the nominal exchange rate Δs, the asset markets return differential $\Delta\tilde{r}$, the relative consumption growth $\Delta\tilde{c}$, the inflation differential $\Delta\hat{p}$, and the change in the real exchange rate Δe. All correlations are based on quarterly observations over the period $1975 : Q1 - 2006 : Q1$.

the core message of this table is related to this separation between correlations with opposite signs for the two components of the real exchange rate. The "asset" component of the real exchange rate - the nominal exchange rate Δs - is positively related to the asset fundamentals, i.e. the asset markets returns differential $\Delta\tilde{r}$, and negatively related to the macro fundamentals, i.e. relative consumption growth $\Delta\tilde{c}$. On the other hand, the "macroeconomic" component of the real exchange rate - the inflation differential $\Delta\hat{p}$ - is positively related to the macro fundamentals, i.e. relative consumption growth, and negatively related to asset fundamentals, i.e. the asset markets returns differential.

Correlations among variables on the same side (asset or macro) seem to be in line with theory. However, variables on one side seem to be disconnected from variables on the other side. For example, the variables on the macro side: relative consumption growth and inflation differential are negatively related to the variables on the asset side: changes in the nominal (and real) exchange rate changes and asset returns differential. Moreover, the nominal exchange rate and the inflation differential are both positively related to the (overall) real exchange rate, but negatively related to each other. In sum, the evidence points towards a dichotomy between the asset and the macroeconomic side of the real exchange rate.

We further investigate the importance of the dichotomy between the two components of the real exchange rate by running time-series regressions on several of these parity conditions. Table 7.5 contains the results from these

Table 7.5: Regression Results

Δs	-0.374 (-0.99)	$\Delta \hat{p}$
0.164** (2.25)		0.257 *** (3.18)
$\Delta \tilde{r}$	-0.628 (-1.27)	$\Delta \tilde{c}$

Note: This table presents results from 4 regressions involving the nominal exchange rate Δs, the asset markets return differential $\Delta \tilde{r}$, the relative consumption growth $\Delta \tilde{c}$, and the inflation differential $\Delta \hat{p}$. The results on each side of the table refer to regression specifications that include the variables in the corresponding corners. All estimations are based on quarterly observations over the period $1975 : Q1 - 2006 : Q1$. The t-statistics, calculated using robust standard errors, and reported in parentheses. Significance at 10%, 5%, and 1% is indicated by *, **, and ***, respectively. The critical values for the t-distribution are 1.645 (10%), 1.960 (5%), and 2.576 (1%).

estimations. Each side of the table refers to estimation results from regression specifications that include the variables in the corresponding corners.

The left side of the table presents results from regressions testing for an equity return parity between UK and USA.[22] This relation is an asset markets counterpart to the uncovered interest rate parity: instead of using risk-free interest rates, this parity uses (expected) stock market returns. The regression yields a significantly positive slope coefficient, indicating that the country with relatively higher returns on its stock market has depreciating currency as well. Hence, this finding gives support to the parity condition on the asset side.[23]

The right side of the table presents results from testing a modified version of the Backus-Smith (1993) condition for efficient international risk-sharing. Instead of using changes in the real exchange rate, this test uses only its macroeconomic component - the inflation differential. The significantly positive coefficient found for this relation is in line with theory and with the findings in chapter 6: consumption growth is higher in the country that experiences a drop in relative prices.[24]

[22] For an elaboration and evidence on the (uncovered) equity return parity see Cappiello and De Santis (2005, 2007).

[23] The results presented here refer to the case of "difference" regressions. When the condition is tested using "level" specification (see chapter 3), the support is even higher.

[24] In fact, the evidence from this regression supports the findings in chapter 6: the Backus-Smith puzzle disappears once the asset component is excluded from the real exchange rate.

Having shown that the parity conditions hold for variables on each side (asset and macro) separately, now we test for the relationships between them. The top side of Table 7.5 presents estimation results for the nominal exchange rate changes - inflation differential relationship, while the bottom side presents the corresponding results for the asset returns differential - relative consumption growth relationship. Both slope coefficient estimates turn negative, though none is significantly different from zero at conventional levels. Although each of these variables is correctly linked to its side of the fundamentals or of the real exchange rate (asset or macro), their relationships with the variables on the other side are with the opposite (anomalous) sign. In turn, these results suggest that there exists a disconnect between the variables on the macroeconomic and the variables on the asset side.

7.3.4 Disconnect and Dichotomy of the Real Exchange Rate

This section synthesized several pieces of evidence about a disconnect and a dichotomy between the macroeconomic and the asset component of the real exchange rate. First, using volatility and correlation figures, it demonstrated that these two components are disconnected from each other. Hence, the asset component appears very volatile compared to macroeconomic fundamentals, while the macroeconomic component appears too smooth compared to asset returns (Table 7.3). Moreover, the correlations of each component with the underlying fundamentals on the other side appear anomalous, with the wrong sign (Tables 7.4 and 7.5). Second, this section presented evidence about a clear dichotomy between the macroeconomic and asset side. Although disconnected from each other, each component of the real exchange rate is correctly correlated/linked to the fundamentals on its own side (Tables 7.4 and 7.5). Finally, though incomparable with each other in volatility terms, both sides imply very similar and plausible values for γ (Table 7.3).

7.4 Conclusion

This study provides a synthesis and some new empirical evidence about the disconnect and the dichotomy that exist between the macroeconomic and the asset side of the real exchange rate. The first part presents empirical evidence on the importance of nominal exchange rate fluctuations for the real exchange rate volatility-disconnect puzzle. Using calculations for implied values of the coefficient of relative risk-aversion, this analysis suggests that both the use of empirical consumption correlations and the stabilization of the nominal exchange rate seem to contribute to the resolution of the volatility-disconnect puzzle. However, while nominal exchange rate stabilization appears to resolve, the complete elimination of the nominal exchange rate seems to over-fix the exchange rate volatility-disconnect puzzle. In fact, using dataset for the Eurozone 12 countries in the period after the introduction of the common currency, the implied value for the coefficient of relative risk-aversion turns very low, even lower than what is implied by logarithmic utility function, for example. In this way, the complete elimination of the nominal exchange rate might point towards a possible anomaly in the opposite direction: once deprived of its asset component, the real exchange rate appears to be very smooth (not very volatile) compared to macroeconomic fundamentals (relative consumption growth in this case).

The second part presents a synthesis of several pieces of evidence about the disconnect and the dichotomy between the macroeconomic and the asset component of the real exchange rate. First, using volatility and correlation figures, it demonstrates that these two components are disconnected from each other. For example, the asset component appears very volatile compared to macroeconomic fundamentals, while the macroeconomic component appears too smooth compared to asset returns. Furthermore, the correlations of each component with the underlying fundamentals on the other side appear anomalous and have the wrong sign. Second, this section presents evidence about a clear dichotomy between the macroeconomic and asset side. Although disconnected from each other, each component of the real exchange rate is correctly correlated/linked to the fundamentals on its own side. In view of several, often opposing, puzzles related to the real exchange rate in the international macro and international finance strands of the literature, an investigation into the underlying reasons for this dichotomy appears promising route for further research.

Chapter 8

Macroeconomic and Financial Aspects of International Risk-Sharing: Concluding Remarks

This chapter presents a recapitulation of the main findings in this thesis. Section 8.1 provides an overview of the main conclusions obtained in each of the empirical chapters. Afterwards, section 8.2 briefly presents an overarching, general conclusion. Finally, section 8.3 discusses the policy relevance of the main conclusions and closes by suggesting several routes for further research.

8.1 Recapitulation per Chapter

The first part of the empirical analysis in chapter 2 replicates some findings from the literature which suggest that international (consumption) risk-sharing is far from perfect. Using data for most developing and industrial countries over the period 1960-2000, the panel analysis suggests that the hypothesis of perfect risk-sharing in consumption is rejected for each group of countries. In fact, the results demonstrate that idiosyncratic consumption growth rates are strongly correlated with idiosyncratic output growth rates, suggesting that only a small portion of total macroeconomic risks is shared across countries. Moreover, and in line with most other empirical studies, this correlation is especially strong for the group of (less-integrated) developing countries, implying very serious departure from the perfect risk-sharing condition.

The second part of the empirical analysis in chapter 2 focuses on the role of workers' remittance flows to developing countries as an alternative channel for international (consumption) risk-sharing. The empirical results suggest that developing countries with above-average workers' remittances inflows per capita also experience relatively higher levels of international risk-sharing in consumption (relatively lower deviations from the perfect risk-sharing condition) over the period 1990-2000. However, though this relationship is true for each subgroup of developing countries (MFIs, LFIs, and transition economies), it is statistically significant only for the group of transition economies. Acknowledging that this effect might be spuriously driven by several data-related problems for this group of countries (especially in the early 1990s), the analysis in chapter 2 concludes that higher than average levels of workers' remittances per capita do not significantly raise international consumption risk-sharing in the developing countries.

The empirical evidence in chapter 3 suggests that the rejection of the UIP condition might be limited to conventional tests that regress the nominal exchange rate change on the interest rate differential. For example, level regressions, which regress the domestic currency return of foreign deposits on the domestic interest rate, generally fail to reject the UIP condition. In fact, these specifications typically produce slope coefficient close to the theoretically-expected value of unity. Moreover, the anchor country interest rate, which can be considered an "omitted" variable in the conventional tests, enters significantly into the baseline specification. Hence, both of these findings lend support to the biasedness hypothesis.

The second part of the empirical analysis investigates the importance of the bias for slope coefficient estimates from conventional tests of UIP. For this purpose, it presents regression results for all possible (bilateral) country-pairs among 10 advanced, industrial countries, and relates these results with the bias (determinants). Although the specific structure of the dataset does not allow for a direct test about the importance of the bias term, the evidence suggests that the main bias determinants are significantly related to the slope coefficient estimates. Finally, chapter 3 concludes with a note of caution: the type of estimation bias identified in the chapter might be important for conventional tests of UIP, but taking it into account does not help in "rehabilitating" the UIP or (completely) solving the UIP puzzle.

Chapter 4 identifies two main inherent inconsistencies of the bilateral SDF approach. First, the discount factors in the bilateral framework are not uniquely determined and crucially depend on the other country included in the risk-sharing calculations. Second, these deviations, i.e. the imprecision in the measurement of marginal utility growth, are most serious for countries with relatively low stock market excess returns (low Sharpe ratios). As an attempt to account for these incoherencies, chapter 4 presents a three-country (trilateral) extension of the SDF model. However, although the trilateral framework demonstrates that the (final) results for the international risk-sharing index are quite robust to the number of countries used in their calculation (risk-sharing indices are even higher in this case), it does not resolve the problems with the inherent inconsistency of the bilateral SDF model.

Chapter 5 is a follow-up to chapter 4. It points out that the SDF-based measures of international risk-sharing crucially depend on the nominal exchange rate behavior/regime. In fact, the empirical analysis (calculation) demonstrates that the SDF approach "mechanically" or "automatically" leads to perfect risk-sharing in episodes of fixed nominal exchange rates. The SDF-based measures imply (literally) perfect risk-sharing for two episodes of fixed nominal exchange rates: first, among all 12 Eurozone countries after the introduction of the common currency; and second, between several emerging markets with fixed nominal exchange rates against the US dollar and the USA. Moreover, this approach implies better integration in terms of higher macroeconomic risk-sharing between several emerging markets (even without completely fixed exchange rates) and the USA than between advanced countries like UK or Japan and the USA. This very counterintuitive result can be considered a major limitation for the use of the SDF approach for cross-country risk-sharing comparisons. In view of the "implausibly" high risk-sharing results obtained for these episodes, chapter 5 concludes that the real exchange rate might be very smooth indeed, but it does not necessarily imply perfect international risk-sharing.

In contrast to the previous two assets markets-based chapters, chapter 6 focuses on a general macroeconomic framework for measuring international risk-sharing. The empirical analysis in this chapter shows that the nominal exchange rate is the main reason for one of the central puzzles in international macroeconomics - the recurrent failure of the efficient international risk-sharing condition, also known as the (relative) consumption-real exchange rate correlation (Backus-Smith) anomaly. Focusing on the 12 Eurozone countries after the introduction of the common currency (time period 1999-2006), the empirical results give strong evidence in support of macroeconomic theory - in the absence of nominal exchange rate changes relative (bilateral) consumption growth is positively related with fluctuations in the real exchange rate. Contrasting these results with several alternative samples with flexible nominal exchange rates (the 12 Eurozone countries before the introduction of the Euro (time period 1986-1999), and 6 advanced, industrial countries over the period 1986-2006), the analysis demonstrates that the nominal exchange rate is the main source of the Backus-Smith puzzle. Though still positively related to the (negative of the) inflation differential in these cases, relative consumption is negatively related to the nominal exchange rate fluctuations. In turn, the chapter concludes with an inquiry about the reasons for the divergent behavior(s) of the macroeconomic (inflation differential) and the asset component (nominal exchange rate) of the real exchange rate in this context.

Beginning with the open question left in chapter 6, the analysis in chapter 7 documents a broad “disconnect” and a “dichotomy” that exist between the macroeconomic and the asset component of the real exchange rate. First, the evidence suggests a clear disconnect: the inflation differential appears too smooth and anomalously correlated to asset markets returns, while the nominal exchange rate appears too volatile and also anomalously correlated to relative consumption. Second, each component seems to be appropriately linked to the fundamentals on its own side: inflation differential is positively related to relative consumption growth (similar as in chapter 6), while the nominal exchange rate is positively related to the asset returns differential (giving support to a risky asset equivalent of the UIP condition tested in chapter 3).

8.2 General Conclusion

The previous chapters provided empirical evidence about the central issue investigated in this thesis - how much international risk-sharing takes place? Alternative approaches, embedded in different theoretical frameworks, suggested divergent results. Most macroeconomic evidence (chapters 2 and 6) suggested that not much risk-sharing takes place across countries. On the other hand, the SDF approach (chapters 4 and 5) suggested almost perfect international risk-sharing. Nevertheless, the discrepancy between the findings obtained using alternative theoretical frameworks necessarily leads to the question: Which of these results are the "correct" ones? How can the evidence from the SDF approach be reconciled with macroeconomic (and portfolio) evidence?

There are two main answers, lines of argumentation: first, the basic international asset pricing condition in chapter 4 does not hold; second, the real exchange rates are sometimes too smooth, but international risk-sharing is not necessarily perfect. The first argument questions the theoretical validity of the equilibrium, no-arbitrage condition between the real exchange rate and the marginal utility growth differential for the corresponding countries. This condition is derived under the strict assumptions that there is free and continuous trade in all assets and there are no pure (zero initial investment) arbitrage opportunities (Chapter 4). Therefore, any departure from these assumptions (transaction and information costs, incompleteness of markets, etc.) might imply a violation of the theoretical asset-pricing condition. The second argument is based on the empirical observation that real exchange rate volatility is (quasi)automatically eliminated with the adoption of a fixed nominal exchange rate (regime) (Chapters 5 and 7). In this case, if real exchange rate changes perfectly measure marginal utility growth differences (as suggested by the SDF approach in chapter 4), then only one policy move (pegging/fixing the nominal exchange rate) is sufficient to eliminate (almost) all macroeconomic risks a country faces (Chapter 5). Therefore, this doubtful (implausible) result raises suspicion about the precision of measuring marginal utility differences (risk-sharing deviations) through real exchange rate changes.

Actually, the potential solution (reconciliation) might be a combination of both arguments. On one hand, chapter 7 (indirectly) provided some evidence in (partial) support of the uncovered equity parity (a risky-asset modified version of the UIP condition discussed in chapter 3). Linking the exchange rate changes with the asset markets returns differential in the corresponding countries, it is similar to the international asset-pricing condition (central in chapters 4 and 5). Although these tests suggested significant relation with the correct (expected) sign, they typically reject the parity hypothesis. Therefore, these findings suggest that the basic international asset-pricing condition might not hold as parity, reflecting the presence of market frictions that are absent from the underlying theoretical framework. On the other hand, the evidence in chapters 5 and 7 demonstrated that fixed/rigid nominal exchange rates lead to very smooth real exchange rates. In turn, if real exchange rate fluctuations really measure deviations from international risk-sharing as suggested by the model in chapter 4, then all countries with fixed/pegged nominal exchange rates should automatically (mechanically) achieve perfect macroeconomic risk-sharing among them. However, this result appears in stark contrast with the general evidence about the risk-sharing environment in these countries. Most of them still have high levels of home bias in equity and bond holdings, enjoy limited access to international capital markets, and lack alternative international insurance mechanisms and/or GDP-indexed securities. In view of all these facts, the perfect risk-sharing figures implied by the SDF model seem very implausible and suggest that real exchange rates can be very smooth, but risk-sharing will not necessarily be perfect. Overall, this discussion suggests that a combination of both arguments might go a long way in bridging the gap between macroeconomic and financial evidence. Finally, note that both of these arguments are in line with the dichotomy in chapter 7.

Irrespective of the numerous differences and discrepancies among them, all approaches reach at least one common conclusion - the (real) exchange rate plays a pivotal role in all measures for international risk-sharing. In fact, in most cases, the exchange rate seems to "drive" the risk-sharing results, largely overshadowing the importance of the co-movement(s) among the underlying consumption or asset returns processes. Therefore, in light of the dichotomy of the real exchange rate argument elaborated in chapter 7, common ground between the different approaches might be sought in a clear and complete separation of the tests conducted on each side of the dichotomy line.

8.3 Policy Relevance and Further Research

What measures are most relevant for policy analysis and recommendations? What figures should the international financial institutions like the IMF follow in evaluating the benefits and in formulating their policy recommendations (e.g. IMF's Integrated Approach; see IMF, 2007) for countries that desire (further) opening of their borders to the global capital market? The balanced discussion above suggests that different (all) types of measures might be useful indicators (for different aspects) in this context. One clear conclusion from this thesis is that all measures should seriously take into account the behavior of the (real) exchange rate. In turn, policy recommendations might rely more on macroeconomic evidence in fixed/rigid nominal exchange rate regimes and probably rely more on the portfolio or asset-pricing measures for (similar) countries with freely floating nominal exchange rate regimes. Inevitably, the right choice or mix between these two options can very often be an art rather than an exact technical procedure.

In the absence of (still) unrealistic macro markets for trading GDP-indexed securities (Shiller, 1993, 2003; Borensztein and Mauro, 2004) and absence of extremely fast integration into the global capital markets, workers remittances might play an increasingly important role in macroeconomic risk-sharing for many (less-integrated) developing countries in the (near) future. The recent surge in workers' remittances flows to developing countries increasingly captures the attention of many international organizations like the World Bank, the IMF, and the OECD, and mobilizes resources for the collection of more accurate data-series. In turn, these developments sound promising for future further research into the macroeconomic (in general) and the consumption-smoothing effects (in particular) of the workers' remittance flows.

Finally, chapter 7 only presented some suggestive, illustrative evidence about the "disconnect" and the "dichotomy" between the two components of the real exchange rate. In light of its probable relation to several puzzles in international macroeconomics, further research into the causes and the (ultimate) consequences of this dichotomy might be fruitful.

Bibliography

[1] Adler, M. and B. Dumas (1983), "International Portfolio Choice and Corporate Finance: A Synthesis", *Journal of Finance*, 38, 925-984.

[2] Aggarwal, R., A. Demirguc-Kunt, and M. S. Martinez Peria (2006), "Do Workers' Remittances Promote Financial Development?", World Bank Policy Research Working Paper, 3957.

[3] Ahearne, A., W. Griever, and F. Warnock (2004), "Information Costs and Home Bias: An Analysis of US Holdings of Foreign Equities", *Journal of International Economics*, 62, 313-336.

[4] Akaike, H. (1973), "Information Theory and an Extension of the Maximum Likelihod Principle", in B. N. Petrov and F. Cszaki (eds.), *Second International Symposium on Information Theory*, Akademiai Kiado, Budapest, 267-281.

[5] Albuquerque, R. (2003), "The Composition of International Capital Flows: Risk Sharing Through Foreign Direct Investment", *Journal of International Economics*, 61, 353-383.

[6] Arrow, K. J. (1971), *Essays in the Theory of Risk-Bearing*, North-Holland, Amsterdam.

[7] Asdrubali, P., B. E. Sorensen, and O. Yosha (1996),"Channels of Interstate Risk-Sharing: United States 1963-1990", *Quarterly Journal of Economics*, 111, 1081-1110.

[8] Backus, D. K., P. J. Kehoe, and F. E. Kydland (1992), "International Real Business Cycles", *Journal of Political Economy*, 100, 745-775.

[9] Backus, D. K., S. Foresi, and C. I. Telmer (1996), "Affine Models of Currency Pricing", NBER Working Paper 5623.

[10] Backus, D. K., S. Foresi, and C. I. Telmer (2001), "Affine Term Structure Models and the Forward Premium Anomaly", *Journal of Finance*, 56, 279-304.

[11] Backus, D. K. and G. W. Smith (1993), "Consumption and Real Exchange Rates in Dynamic Economies with Non-traded Goods", *Journal of International Economics*, 35, 297-316.

[12] Baillie, R.T. and T. Bollerslev (2000), "The Forward Premium Puzzle is Not as Bad as You Think", *Journal of International Money and Finance*, 19, 471-488.

[13] Bansal, R. and M. Dahlquist (2000), "The Forward Premium Puzzle, Different Tales from Developed and Emerging Economies", *Journal of International Economics*, 51, 115-144.

[14] Baxter, M. (1995), "International Trade and Business Cycles" , in G. H. Grossman and K. Rogoff (eds.), *Handbook of International Economics*, 3, 1801-1864.

[15] Baxter, M. and M. J. Crucini (1995), "Business Cycles and the Asset Structure of Foreign Trade", *International Economic Review*, 36, 821-854.

[16] Baxter, M. and U. J. Jermann (1997),"The International Diversification Puzzle is Worse Than You Think", *American Economic Review*, 87, 170-180.

[17] Baxter, M. and A. C. Stockman (1989), "Business Cycles and the Exchange Rate Regime: Some International Evidence", *Journal of Monetary Economics*, 23, 377-400.

[18] Bekaert, G. and M. S. Urias (1996), "Diversification, Integration, and Emerging Market Closed-End Funds", *Journal of Finance*, 51, 835-869.

[19] Bekaert, G., M. Wei, and Y. Xing (2006), "Uncovered Interest Rate Parity and the Term Structure", *Journal of Internationa Money and Finance*, 26, 1038-1069.

[20] Benigno, G. and C. Thoenissen (2005), "On the Consumption-Real Exchange Rate Anomaly", Bank of England Working Paper 254.

[21] Bodenstein, M. (2005), "International Asset Markets and Real Exchange Rate Volatility", mimeo, Northwestern University.

[22] Bordo, M.D. and R. MacDonald (2003), "The Inter-War Gold Exchange Standard: Credibility and Monetary Independence", *Journal of International Money and Finance*, 22, 1-32.

[23] Borensztein, E., M. Chamon, O. Jeanne, P. Mauro, and J. Zettelmeyer (2004),"Sovereign Debt Structure for Crisis Prevention", IMF Occasional Paper 237.

[24] Borensztein, E. and P. Mauro (2004), "The Case for GDP-Indexed Bonds", *Economic Policy*, 19, 165-216.

[25] Brandt, M. W., J. H. Cochrane, and P. Santa-Clara (2006), "International Risk Sharing is Better Than You Think, or Exchange Rates are Too Smooth", *Journal of Monetary Economics*, 53, 671-698.

[26] Brandt, M. W. and P. Santa-Clara (2002), "Simulated Likelihood Estimation of Diffusions with an Application to Exchange Rate Dynamics in Incomplete Markets", *Journal of Financial Economics*, 63, 161-210.

[27] Brennan, M. and H. Cao (1997), "International Portfolio Investment Flows", *Journal of Finance*, 52, 1851-1880.

[28] Buch, C. M., and A. Kuckulenz (2004),"Worker Remittances and Capital Flows to Developing Countries", ZEW Discussion Paper 04-31.

[29] Bugamelli, M. and F. Paterno (2005), "Do Workers' Remittances Reduce the Probability of Current Account Reversals?", World Bank Policy Research Working Paper 3766.

[30] Cappiello, L. and R. A. De Santis (2005), ‘Explaining Exchange Rate Dynamics: The Uncovered Equity Return Parity Condition”, ECB Working Paper 529.

[31] Cappiello, L. and R. A. De Santis (2007), “The Uncovered Return Parity Condition”, ECB Working Paper 812.

[32] Chaboud, A.P. and J.H. Wright (2005), “Uncovered Interest Parity, It Works, but not for Long”, *Journal of International Economics*, 66, 349-362.

[33] Chamon, M. and P. Mauro (2005), “Pricing Growth-Indexed Bonds”, IMF Working Paper 05/216.

[34] Chari, V. V., P. J. Kehoe, and E. R. McGrattan (1997), “Monetary Shocks and Real Exchange Rates in Sticky Price Models of International Business Cycles”, NBER Working Paper 5876.

[35] Chari, V. V., P. J. Kehoe, and E. R. McGrattan (2002), “Can Sticky Price Models Generate Volatile and Persistent Real Exchange Rates?”, *Review of Economic Studies*, 69, 633-663.

[36] Chinn, M.D. (2006), “The (Partial) Rehabilitation of Interest Rate Parity in the Floating Rate Era: Longer Horizons, Alternative Expectations, and Emerging Markets”, *Journal of International Money and Finance*, 25, 7-21.

[37] Chinn, M.D. and J.A. Frankel (2002), “Survey Data on Exchange Rate Expectations: More Currencies, More Horizons, More Tests”, in W. Allen and D. Dickinson, *Monetary Policy, Capital Flows, and Financial Market Developments in an Era of Financial Globalization: Essays in Honour of Max Fry*, Routledge, London.

[38] Chinn, M.D. and G. Meredith (2005), Testing Uncovered Interest Parity at Short and Long Horizons During the Post-Bretton Woods Era, NBER Working Paper 11077.

[39] Cochrane, J. H. (1991), “A Simple Test Of Consumption Insurance”, *Journal of Political Economy*, 99, 957-976.

[40] Cochrane, J. H. (2004), *Asset Pricing*, Princeton University Press, Princeton.

[41] Cole, H. and M. Obstfeld (1991), "Commodity Trade and International Risk Sharing: How Much Do Financial Markets Matter?",*Journal of Monetary Economics*, 28, 3-24.

[42] Cooper, I. A. and E. Kaplanis (1994), "Home Bias in Equity Portfolios, Inflation Hedging, and International Capital Market Equilibrium", *Review of Financial Studies*, 7, 45-60.

[43] Corsetti, G., L. Dedola, and S. Leduc (2004), "International Risk-Sharing and the Transmission of Productivity Shocks", ECB Working Paper Series 308.

[44] Corsetti, G., L. Dedola, and S. Leduc (2007), "International Risk-Sharing and the Transmission of Productivity Shocks", *Review of Economic Studies* (forthcoming).

[45] Davis, S. J., J. Nalewaik, and P. Willen (2000), "On the Gains from International Trade in Risky Financial Assets", NBER Working Paper 7796.

[46] Dickey, D. A. and W. A. Fuller (1979), "Distribution of the Estimators for Autoregressive Time Series with A Unit Root", *Journal of the American Statistical Association*, 74, 427-431.

[47] Dixon, H. (1999), "Controversy: Exchange Rates and Fundamentals", *Economic Journal*, 109, F652-F654.

[48] Engel, Ch. (2000), "Comments on Obstfeld and Rogoff's 'Six Major Puzzles in International Macroeconomics: Is There a Common Cause?"', NBER Working Paper 7818.

[49] Engel, C. (2000), "PPP May Not Hold After All", *Journal of International Economics*, 51, 243-273.

[50] Engel, C. and J. H. Rogers (1996), "How Wide is the Border", *American Economic Review*, 86, 1112-1125.

[51] Errunza, V., K. Hogan, and M.-W. Hung (1999), "Can the Gains from International Diversification Be Achieved without Trading Abroad?", *Journal of Finance*, 54, 2075-2107.

[52] Fama, E.F. (1984), "Forward and Spot Exchange Rates", *Journal of Monetary Economics*, 14, 319-338.

[53] Feldstein, M. and C. Horioka (1980), "Domestic Savings and International Capital Flows", *Economic Journal*, 90, 314-329.

[54] Flood, R.P. and A.K. Rose (1996), "Fixes: Of the Forward Discount Puzzle", *Review of Economics and Statistics*, 78, 748-752.

[55] Flood, R.P. and A.K. Rose (1999), "Understanding Exchange Rate Volatility Without the Contrivance of Macroeconomics", *Economic Journal*, 109, $F660$-$F672$.

[56] Flandreau, M. (2005), "Home Biases, 19th Century Type", CEPR Discussion Paper 5398.

[57] Frankel, J.A. and K.A. Froot (1987), "Using Survey Data to Test Standard Propositions Regarding Exchange Rate Expectations", *American Economic Review*, 77, 133-153.

[58] Frankel, J.A. and J. Poonawala (2004), "The Forward Market in Emerging Currencies, Less Biased Than in Major Currencies", mimeo, Kennedy School of Government, Harvard University.

[59] French, K. R. and J. M. Poterba (1991), "International Diversification and International Equity Markets", *American Economic Review*, 81, 222-226.

[60] Friend, I. and M. E. Blume (1975), "The Demand for Risky Assets", *American Economic Review*, 65, 900-922.

[61] Froot, K. A. and K. Rogoff (1995), "Perspectives on PPP and Long-Run Real Exchange Rates", in G. M. Grossman and K. Rogoff (eds.), *Handbook of International Economics*, 3, North-Holland.

[62] Froot, K.A. and R.H. Thaler (1990), “Anomalies: Foreign Exchange”, *Journal of Economic Perspectives*, 4, 179-192.

[63] Giuliano, P. and M. Ruiz-Arranz (2006), “Remittances, Financial Development, and Growth”, IZA Discussion Paper 2160.

[64] Glick, R. and K. Rogoff (1995), “Global v Country-Specific Productivity Shocks and the Current Account”, *Journal of Monetary Economics*, 35, 159-192.

[65] Gorman, L. R. and B. N. Jorgensen (1996), “Domestic Versus International Portfolio Selections: A Statistical Examination of the Home Bias”, Working Paper, Kellogg Graduate School of Management.

[66] Hansen, L. P. and R. Jagannathan (1991), “Implications of Security Market Data for Models of Dynamic Economies”, *Journal of Political Economy*, 99, 225-262.

[67] Hausman, J. A. (1978), “Specification Tests in Econometrics”, *Econometrica*, 46, 1251-1271.

[68] Haynes, S. E. and J. A. Stone (1981), “On the Mark: Comment”, *American Economic Review*, 71, 1060-1067.

[69] Haynes, S. E. and J. A. Stone (1982), “Spurious Tests and Sign Reversals in International Economics”, *Southern Economic Journal*, 48, 868-876.

[70] Hess, G. D. and K. Shin (2006), “Understanding the Backus-Smith Puzzle: It’s the (Nominal) Exchange Rate, Stupid”, MPRA Paper 696.

[71] Heston, S. L. and K. G. Roewenhorst (1994), “Does Industrial Structure Explain the Benefits of International Diversification?”, *Journal of Financial Economics*, 36, 3-27.

[72] Heston, A., R. Summers, and B. Aten (2000), Penn World Table Version 6.1, Center for International Comparisons at the University of Pennsylvania (CICUP).

[73] Huisman, R., K.G. Koedijk, C.J.M. Kool, and F. Nissen (1998), "Extreme Support for Uncovered Interest Parity", *Journal of International Money and Finance*, 17, 211-228.

[74] Imbs, J. and P. Mauro (2007), "Pooling Risk Among Countries", IMF Working Paper 07/132.

[75] IMF (1993), *Balance of Payments Manual*, 5th edition, Washington.

[76] IMF (2004), *Balance of Payments Statistics Yearbook*, Washington.

[77] IMF (2005), "Two Current Issues Facing Developing Countries", *World Economic Outlook*, Chapter II, April 2005.

[78] IMF (2005), "Globalization and External Imbalances" Chapter III, *World Economic Outlook*, April 2005.

[79] IMF (2007), "Reaping the Benefits of Financial Globalization", IMF Discussion Paper, June 2007.

[80] Isard, P. (2006), "Uncovered Interest Parity", IMF Working Paper 06/96.

[81] Kandel, S. and R. F. Stambaugh (1991), "Asset Returns and Intertemporal Preferences", *Journal of Monetary Economics*, 27, 39-71.

[82] Kang, J. S. (2006), "Consumption and Real Exchange Rates in An Economy with Limited Commitment", mimeo, University of Wisconsin at Madison.

[83] Kang, J. S. (2007), "Consumption and Real Exchange Rates in An Economy with Private Information", mimeo, University of Wisconsin at Madison.

[84] Kang, J.-K. and R. M. Stulz (1997), "Why is There Home Bias? An Analysis of Foreign Portfolio Equity Ownership in Japan", *Journal of Financial Economics*, 46, 3-28.

[85] Kocherlakota, N. R. (1996), "The Equity Premium: It's Still a Puzzle", *Journal of Economic Literature*, 34, 42-71.

[86] Kollmann, R. (1995), "Consumption, Real Exchange Rates, and the Structure of International Capital Markets", *Journal of International Money and Finance*, 14, 191-211.

[87] Kool, C.J.M. and D.L. Thornton (2004), "A Note on the Expectations Hypothesis at the Founding of the Fed", *Journal of Banking and Finance*, 28, 3055-3068.

[88] Kose, M. A., E. S. Prasad, and M. E. Terrones (2007), "How Does Financial Globalization Affect Risk-Sharing? Patterns and Channels", IMF Working Paper 07/238

[89] Kwiatkowski, D., P. C. B. Phillips, P. Schmidt, and Y. Shin (1992), "Testing the Null Hypothesis of Stationarity Against the Alternative of A Unit Root: How Sure Are We That Economic Time Series Have A Unit Root?", *Journal of Econometrics*, 54, 159-178.

[90] Kydland, F. E. and E. C. Prescott (1982), "Time to Build and Aggregate Fluctuations", *Econometrica*, 50, 1345-1370.

[91] Lane, Ph. R. and G. M. Milesi-Ferretti (2001), "The External Wealth of Nations: Measures of Foreign Assets and Liabilities for Industrial and Developing Contries", *Journal of International Economics*, 55, 263-294.

[92] Lane, Ph. R. and G. M. Milesi-Ferretti (2003), "International Financial Integration", *IMF Staff Papers*, 50, 82-113.

[93] Lane, Ph. R. and G. M. Milesi-Ferretti (2005), "Financial Globalization and Exchange Rates", IMF Working Paper 05/03.

[94] Lane, Ph. R. and G. M. Milesi-Ferretti (2006), "The External Wealth of Nations Mark II: Revised and Extended Estimates of Foreign Assets and Liabilities, 1970-2004", IMF Working Paper 06/69.

[95] Lewis, K. K. (1996), "What Can Explain the Apparent Lack of International Consumption Risk Sharing?", *Journal of Political Economy*, 104, 267-297.

[96] Lewis, K. K. (1999), "Trying to Explain Home Bias in Equities and Consumption", *Journal of Economic Literature*, 37, 571-608.

[97] Lewis, K. K. (2000), "Why Do Stocks and Consumption Imply Such Different Gains from International Risk-Sharing", *Journal of International Economics*, 52, 1-35.

[98] Lothian, J.R. and C.H. McCarthy (2002), "Real Exchange Rate Behavior under Fixed and Floating Exchange Rate Regimes", *Manchester School*, 70, 229-245.

[99] Lothian, J.R. and L. Wu (2003), Uncovered Interest-Rate Parity over the Past Two Centuries, mimeo, Fordham University.

[100] Lucas, R. E. Jr. (1987), *Models of Business Cycles*, Oxford and New York, Blackwell.

[101] de Luna Martinez, J. (2005), "Workers' Remittances to Developing Countries: A Survey with Central Banks on Selected Public Policy Issues", World Bank Policy Research Working Paper No. 3638.

[102] Lustig, H. and A. Verdelhan (2007), "The Cross Section of Foreign Currency Risk Premia and Consumption Growth Risk", *American Economic Review*, 97, 89-117.

[103] MacDonald, R. (1999), "Exchange Rate Behaviour: Are Fundamentals Important?", *Economic Journal*, 109, F673-F691.

[104] Mace, B. (1991), "Full Insurance in the Presence of Aggregate Uncertainty", *Journal of Political Economy*, 99, 928-956.

[105] Maddala, G.S. and S. Wu (1999), "A Comparative Study of Unit-Root Tests with Panel Data and A New Simple Test", *Oxford Bulletin of Economics and Statistics*, 61, 631-652.

[106] Mankiw, N. G. and S. P. Zeldes (1991), "The Consumption of Stockholders and Non-Stockholders", *Journal of Financial Economics*, 29, 97-112.

[107] Mark, N. C. (2001), *Internationa Macroeconomics and Finance: Theory and Econometric Methods*, Blackwell Publishing.

[108] Mehra, R. and E. C. Prescott (1985), "The Equity Premium: A Puzzle", *Journal of Monetary Economics*, 15, 145-161.

[109] Mehra, R. and E. C. Prescott (1988), "The Equity Risk Premium: A Solution?", *Journal of Monetary Economics*, 22, 133-136.

[110] Mussa, M. (1986), "Nominal Exchange Rate Regimes and the Behavior of Real Exchange Rates: Evidence and Implications", *Carnegie-Rochester Conference Series on Public Policy*, 25, 117-214.

[111] Newey, W. K. and K. D. West (1987), "A Simple, Positive Semi-Definite, Heteroskedasticity and Autocorrelation Consistent Covariance Matrix", *Econometrica*, 55, 703-708.

[112] Obstfeld, M. (1994), "Risk-Taking, Global Diversification, and Growth", *American Economic Review*, 84, 1310-1329.

[113] Obstfeld, M. (1998), "The Global Capital Market: Benefactor or Menace?", *Journal of Economic Perspectives*, 12, 9-30.

[114] Obstfeld, M. and K. Rogoff (1995), "The Intertemporal Approach to the Current Account" , in G. H. Grossman and K. Rogoff (eds.), *Handbook of International Economics*, 3, 1731-1799.

[115] Obstfeld, M. and K. Rogoff (1996), *Foundatons of International Macroeconomics*, MIT Press.

[116] Obstfeld, M. and K. Rogoff (2001), "The Six Major Puzzles in International Macroeconomics: Is There A Common Cause?", in B. Bernanke and K. Rogoff (2001), *NBER Macroeconomics Annual 2000*, 339-390.

[117] Pakko, M. R. (1998), "Characterizing Cross-Country Consumption Correlations", *Review of Economics and Statistics*, 80, 169-174.

[118] Pakko, M. R. (2004), "A Spectral Analysis of the Cross-Country Consumption Correlation Puzzle", *Economics Letters*, 84, 341-347.

[119] Prasad, E., K. Rogoff, S.-J. Wei, and M.A. Kose (2003), *Effects of Globalization on Developing Countries: Some Empirical Evidence*, IMF Occasional Paper 220.

[120] Ratha, D. (2003), "Workers' Remittances: An Important and Stable Source of External Development Finance", in *Global Development Finance: Striving for Stability in Development Finance*, World Bank, 157-175.

[121] Ravn, M. O. (2001), "Consumption Dynamics and Real Exchange Rates", CEPR Discussion Paper 2940.

[122] Reinke, J. and N. Patterson (2005), "Remittances in the Balance of Payments Framework", *mimeo*, International Technical Meeting on Measuring Remittances, World Bank and IMF, Washington.

[123] Rogoff, K. (1996), "The Purchasing Power Parity Puzzle", *Journal of Economic Literature*, 34, 647-668.

[124] Rowland, P. F. and L. L. Tesar (2004), "Multinationals and the Gains from International Diversification", *Review of Economic Dynamics*, 7, 789-826.

[125] Sachs, J. (1982), "The Current Account in the Macroeconomic Adjustment Process" , *Scandinavian Journal of Economics*, 84, 147-159.

[126] Schwarz, G. (1978), "Estimating the Dimension of a Model", *Annals of Statistics*, 6, 461-464.

[127] Schwert, G. W. (1989), "Tests for Unit Roots: A Monte Carlo Investigation", *Journal of Business and Economic Statistics*, 7, 147-160.

[128] Shiller, R. J. (1993), *Macro Markets: Creating Institutions for Managing Society's Largest Economic Risks*, Clarendon Press, Oxford, UK.

[129] Shiller, R. J. (2003), *The New Financial Order: Risk in the* 21*st Century*, Princeton University Press, Princeton, NJ, USA.

[130] Sill, K. (2001), "The Gains from International Risk-Sharing", *Business Review*, Q3, 23-32.

[131] Sorensen, B. E. and O. Yosha (1998), "International Risk-Sharing and European Monetary Unification", *Journal of International Economics*, 45, 211-238.

[132] Sorensen, B. E., Y.-T. Wu, O. Yosha, and Y. Zhu (2005), "Home Bias and International Risk-Sharing: Twin Puzzles Separated at Birth", CEPR Discussion Paper 5113.

[133] Sorensen, B. E., Y.-T. Wu, O. Yosha, and Y. Zhu (2007), "Home Bias and International Risk Sharing: Twin Puzzles Separated at Birth", *Journal of International Money and Finance*, 26, 587-605.

[134] Stulz, R. M. (1981), "A Model of International Asset Pricing", *Journal of Financial Economics*, 9, 383-406.

[135] Tesar, L. and I. Werner (1995), "Home Bias and High Turnover", *Journal of International Money and Finance*, 14, 467-492.

[136] Thoenissen, C. (2006), "Real Exchange Rate Volatility and Asset Market Structure", Centre for Dynamic Macroeconomic Analysis Working Paper 06/09.

[137] Thornton, D.L. (2005), "Tests of the Expectations Hypothesis: Resolving the Anomalies when the Short-Term Rate is the Federal Funds Rate", *Journal of Banking and Finance*, 29, 2541-2556.

[138] Thornton, D.L. (2006), "Tests of the Expectations Hypothesis: Resolving the Campbell-Shiller Paradox", *Journal of Money, Credit, and Banking*, 38, 511-542.

[139] Van Wincoop, E. (1994), "Welfare Gains from International Risksharing", *Journal of Monetary Economics*, 34, 175-200.

[140] Van Wincoop, E. (1999), "How Big Are Potential Welfare Gains from International Risk-Sharing?", *Journal of International Economics*, 47, 109-135.

[141] Verbeek, M. (2004), *A Guide to Modern Econometrics* (2nd ed.), John Wiley and Sons, West Sussex, England.

Samenvatting

Internationale risicodeling is een van de meest belangrijke voordelen van internationale financiële integratie, en krijgt daardoor steeds meer aandacht. Zolang verschillende nationale economieën niet perfect gecorreleerd zijn, zijn er mogelijkheden voor risicodeling middels de handel in activa en goederen. De inwoners van deze economieën kunnen het idiosyncratische, landspecifieke deel van het risico dat zij lopen weg diversifiëren in een grotere, wereldwijde (kapitaal)markt. Hierdoor kunnen zij additionele welvaartswinsten behalen, die anders binnen de landsgrenzen buiten bereik blijven. Wanneer eenieder optimaal gebruik maakt van deze voordelen van international risicodeling, zou in principe het idiosyncratische (landspecifieke) risico compleet verdwijnen, zodat ieder land alleen haar (proportionele deel van het) globale systeemrisico draagt. Hierdoor, zou het gezamenlijk, niet diversifieerbare risico wereldwijd het zelfde effect moeten hebben, en de groeivoeten van het marginale nut zouden overal gelijk moeten zijn.

Dezelfde voordelen van een grotere en diepere kapitaalmarkt zijn veelvuldig gebruikt door internationale financiële instellingen om ontwikkelingslanden ertoe aan te zetten hun kapitaalmarkten meer met de geïndustrialiseerde wereld te integreren. Het is vanzelfsprekend te verwachten dat toekomstige vooruitgang op het gebied van financiële globalisatie af zal hangen van de in het verleden daadwerkelijk gerealiseerde welvaartsvoordelen. Daarom is het voor een evaluatie van het proces van financiële integratie van groot belang om over goede maatstaven voor de mate van internationale risicodeling te beschikken.

Nochtans worden in de academische literatuur verschillende methodes gesuggereerd, die tot afwijkende conclusies met betrekking tot de mate risicodeling tussen landen. Zo wijzen macro-economische methodes, gebaseerd op con-

sumptie en productie data, op beperkte risicodeling. Dit resultaat wordt verder geïllustreerd aan de hand van het feit dat consumptie groeivoeten in de OECD landen slecht zeer zwak gecorreleerd zijn, vaak nog zwakker dan productie groeivoeten. Ook is in vele opkomende economieën, na de toetreding to de wereldwijde kapitaalmarkt in de jaren '90, de volatiliteit van de consumptie toegenomen, vergeleken met de volatiliteit van de productie. Tot slot hangt de relatieve consumptiegroei meestal negatief af van de veranderingen in de reële wisselkoers.

Hier staat tegenover, dat prijsmodellen voor activa, gebaseerd op surplus rendementen, een zeer hoge mate van internationale risicodeling suggereren. Volgens deze modellen zijn reële wisselkoersen, als maatstaf voor het verschil in de groeivoeten van het marginale nut, veel minder volatiel dan stochastische discontovoeten, die als maatstaf voor de marginale groeivoeten in elk land fungeren.

Daarom is het, in het licht van de te verwachten groei van het aantal ontwikkelingslanden dat aan het financiële integratie proces zal deelnemen, van belang om deze verschillende methodes verder te analyseren, ten einde tot een overtuigende raamwerk voor het kwantificeren van internationale risicodeling te komen.

Het doel van dit proefschrift is om de mate van risicodeling tussen landen zoals die resulteert uit het gebruik van zowel macro-economische als financiële methodes, in één compact, synthetisch raamwerk te vatten. In deze synthese functioneert de reële wisselkoers als een natuurlijke geleider, die beide methodes met elkaar verbindt.

Om de centrale vraag - hoeveel internationale risicodeling vindt er daadwerkelijk plaats - te beantwoorden, worden in dit proefschrift verscheidene, soms met elkaar vervlochten onderzoeksvelden in de macro-economische en financiële literatuur gebruikt. De 'kern' van de literatuur waarop dit proefschrift zich baseert, bestaat uit drie delen: macro-economische conjunctuurmodellen, portfolio theorie, en internationale prijsmodellen voor activa, gebaseerd op het afwezig zijn van arbitrage mogelijkheden. Daarnaast raakt dit proefschrift nog aan een aantal andere delen van de literatuur: anderen kanalen waarlangs risicodeling plaatsvindt (in het bijzonder internationale overschrijvingen door werknemers), testen van internationale pariteitcondities voor aandelenrendementen en rentes, en studies met betrekking tot de gebrekkige relatie tussen reële wisselkoersen en fundamenten en de (niet-)neutraliteit van nominale wisselkoersregimes.

Hoofdstuk 2 behandelt alternatieve kanalen voor internationale risicodeling. Het eerste deel van de empirische analyse in dit hoofdstuk repliceert een aantal bevindingen uit de literatuur, die suggereren dat internationale (consumptie) risicodeling verre van perfect is. In een panel analyse aan de hand van data voor ontwikkelingslanden en geïndustrialiseerde landen voor de periode 1960-2000 wordt de hypothese dat er bij consumptie sprake is van perfect risicodeling verworpen voor beide groepen landen. Uit de resultaten blijkt zelfs dat de idiosyncratische groeivoeten van consumptie sterk gecorreleerd zijn met de idiosyncratische groeivoeten van productie, hetgeen suggereert dat slechts een klein deel van het totale macro-economische risico daadwerkelijk gedeeld wordt tussen landen. Bovendien, en in overeenstemming met andere empirische studies, blijkt deze correlatie vooral voor de groep (minder geïntegreerde) ontwikkelingslanden erg sterk te zijn, hetgeen impliceert dat vooral deze landen ver van perfecte risicodeling verwijderd zijn.

Het tweede deel van de empirische analyse in hoofdstuk 2 concentreert zich op de rol van internationale overschrijvingen door werknemers naar hun moederland als een alternatief kanaal voor internationale risicodeling (van consumptie). Panel schattingsresultaten voor de periode 1990-2000 suggereren dat ontwikkelingslanden met bovengemiddelde internationale overschrijvingen (per capita) tevens gekenmerkt worden door een hogere mate van international risicodeling van consumptie. Echter, hoewel dit resultaat opgang doet voor elke subgroep van ontwikkelingslanden (meer financieel geïntegreerd (MFI), minder financieel geïntegreerd (LFI) en transitielanden), is het alleen voor de groep transitielanden statistisch significant. Aangezien dit laatste resultaat het gevolg kan zijn van een aantal data problemen voor deze groep landen (vooral in het begin van de jaren '90), concludeert hoofdstuk 2 dat bovengemiddelde internationale overschrijvingen door werknemers niet significant bijdragen aan de internationale risicodeling van consumptie in ontwikkelingslanden.

Hoofdstuk 3 gaat in op de mogelijkheid dat empirische testen van de ongedekte rentepariteit (UIP) mogelijk vertekend zijn. Deze pariteitconditie stelt dat de verwachte rendementen op twee vergelijkbare (risicovrije) activa uitgedrukt in verschillende munteenheden gelijk moet zijn. Bijgevolg wordt deze renteparitei een van de meest gebruikte maatstaven voor de intensiteit van het proces van internationale financiële integratie. Hoewel op de theoretische onderbouwing van UIP niets aan te merken is, wordt UIP in empirische testen consequent verworpen. Het begin van dit hoofdstuk toont aan dat de standaard schattingsmethode tot vertekening kan leiden indien het onderliggende proces waarbij de schattingsgegevens gegenereerd worden afwijkt van hetgeen theoretisch verondersteld wordt.

De empirische resultaten in hoofdstuk 3 suggereren dat UIP vooral verworpen wordt in conventionele testen, waarbij een regressie uitgevoerd wordt van de verandering in de nominale wisselkoers op het renteverschil. Wanneer bijvoorbeeld een regressie uitgevoerd wordt van het rendement op een buitenlands deposito (uitgedrukt in de binnenlandse munteenheid) op de binnenlandse rentevoet, kan UIP meestal niet verworpen worden. Het schatten van een dergelijke specificatie resulteert meestal in een coëfficiënt die zeer dicht bij de theoretisch te verwachte waarde van één ligt. Bovendien is de rentevoet (van het ankerland), die als een 'weggelaten' variabele beschouwd kan worden in conventionele testen, significant in deze specificatie. Deze resultaten bevestigen dan ook de hypothese dat standaard testen vertekend zijn.

Het resterende deel van de empirische analyse in hoofdstuk 3 gaat nader in op het belang van deze vertekening bij conventionele UIP testen. Daartoe worden voor 10 geavanceerde, industriële landen (bilaterale) schattingsresultaten gepresenteerd voor alle mogelijke landenparen. Deze resultaten worden vervolgens gerelateerd aan variabelen die mogelijke de vertekening verklaren. Hoewel de structuur van de dataset een direct test van het belang van de vertekening niet toestaat, verklaren de variabelen die vertekening meten een significant deel van de UIP coëfficiënt. Het hoofdstuk eindigt met een waarschuwing: hoewel de vertekening die hier beschreven is van belang kan zijn voor conventionele testen van UIP, is het helaas niet zo dat het controleren voor deze vertekening voor een volledige 'rehabilitatie' van UIP of tot een oplossing van de UIP puzzel leidt.

Hoofdstukken 4 en 5 concentreren zich op het meten van internationale risicodeling met behulp van de stochastische discontovoet methode (SDF). Deze methode suggereert een hoge mate van diversificatie van risico tussen landen. Daarmee zijn deze resultaten in tegenspraak met de meeste macro-economische studies die gebaseerd zijn op internationale reële conjunctuurcyclus modellen en financiële studies die op portfolio theorie gebaseerd zijn. De (bilaterale) SDF methode begint met de standaard (internationale) voorwaarde voor het prijzen van activa, die stelt dat veranderingen in de reële wisselkoers exact overeen dienen te komen met verschillen in veranderingen in de stochastische discontovoeten voor twee landen. De resultaten laten zien dat een aanzienlijke hoeveelheid macro-economische risico kan worden gedeeld tussen landen, aangezien de discontovoeten erg volatiel zijn. Ook blijkt dat een flink deel van dit macro-economisch risico daadwerkelijk te worden gedeeld aangezien reële wisselkoersen veel minder volatiel zijn. Hoofdstukken 4 en 5 presenteren een aantal beperkingen en extensies van deze aanpak.

Hoofdstuk 4 presenteert het theoretische raamwerk voor de (bilaterale) SDF methode en repliceert een aantal eerdere resultaten uit de literatuur. Bovendien identificeert het twee belangrijke inconsistenties in de bilaterale SDF methode. Ten eerste zijn de discontovoeten in het bilaterale raamwerk niet uniek bepaald, aangezien ze cruciaal afhangen van het andere land wanneer de risicodeling berekend wordt. Ten tweede, is onnauwkeurigheid waarmee de groei van het marginale nut gemeten wordt vooral groot voor landen met relatief lage surplusrendementen op de aandelenmarkt (lage Sharpe ratios). In een poging voor deze inconsistenties te controleren, presenteert hoofdstuk 4 drie landen (trilaterale) extensie van het SDF model. Echter, hoewel het trilaterale raamwerk aantoont dat de uiteindelijke index van international risicodeling vrij robuust is voor het aantal landen waarop de berekening gebaseerd is, wordt de inherente inconsistentie van het bilaterale SDF model niet opgelost.

Hoofdstuk 5 is een vervolg op hoofdstuk 4, en het laat zien dat de SDF maatstaven voor internationale risicodeling sterk afhangen van het nominale wisselkoers gedrag/regime. De empirische analyse (berekening) laat zien dat de SDF methode 'mechanisch' of 'automatisch' leidt tot perfecte risicodeling gedurende periodes waarin de nominale wisselkoers vastligt. De op SDF gebaseerde maatstaven duiden (letterlijk) op perfecte risicodeling gedurende twee periodes waarin nominale wisselkoersen vastlagen: ten eerste, voor de

12 Eurozone landen na de introductie van de gemeenschappelijke munt; ten tweede, voor een aantal opkomende economieën die hun nominale wisselkoers met de Amerikaanse dollar vastgelegd hebben. Bovendien, laat deze methode zien dat er meer internationale macro-economische risicodeling plaatsvindt tussen een aantal opkomende economieën (zelfs zonder volledig vastliggende wisselkoersen) en de Verenigde Staten dan tussen geavanceerde landen zoals het Verenigd Koninkrijk of Japan en de Verenigde Staten. Dit bijzonder contra-intuïtieve resultaat toont de beperkingen van de SDF methode voor het vergelijken van international risicodeling aan. Hoofdstuk 5 concludeert dan ook zelfs een erg gelijkmatige reële wisselkoers niet noodzakelijk impliceert dat van perfecte risicodeling sprake is.

In tegenstelling tot de voorgaande twee hoofdstukken die op activa markten gebaseerd zijn, concentreert hoofdstuk 6 zich op een algemeen macro-economisch raamwerk voor het meten van internationale risicodeling. Dit hoofdstuk onderzoekt het belang van nominale wisselkoers fluctuaties voor de onregelmatige correlatie tussen de relatieve consumptiegroei en reële wisselkoers veranderingen. Deze onregelmatige relatie gaat in tegen de voorwaardes voor efficiënte international risicodeling (van consumptie) zoals die in internationale conjunctuurmodellen afgeleid worden. Deze modellen suggereren dat consumptie relatief hoog zou moeten zijn in landen waar de relatieve prijzen dalen. Deze consumptie reële wisselkoers correlatie anomalie vormt dan ook een van de belangrijkste puzzels in de internationale macro economie, en staat bekend als de Backus-Smith puzzel. Het belangrijkste doel van dit hoofdstuk, is te onderzoeken in welke mate deze anomalie het gevolg is van nominale wisselkoersveranderingen. Met het verdwijnen van de intra-Eurozone nominale wisselkoers na de introductie van de Euro ontstond een 'natuurlijke' test voor deze hypothese.

Voor de 12 Eurozone landen na de introductie van de gemeenschappelijke munt (in de period 1999-2006) laat de empirische analyse zien dat de nominale wisselkoers inderdaad de voornaamste verklaring voor de puzzel vormt - bij afwezigheid van nominale wisselkoers fluctuaties hangt de relatieve (bilaterale) consumptiegroei positief samen met reële wisselkoers fluctuaties, zoals de macro-economische theorie (RBC modellen) voorspelt. Wanneer deze resultaten worden vergeleken met een aantal samples met flexibele wisselkoersen (de 12 Eurozone landen voor de introductie van de Euro (in de periode 1986-1999), en 6 geavanceerde, industriële landen gedurende de periode 1986-2000), blijkt dat de nominale wisselkoers de belangrijkste verklarende factor voor de Backus-Smith puzzel is. Hoewel relatieve consumptie nog steeds positief samenhangt met het (omgekeerde) inflatieverschil, hangt het negatief samen met de nominale wisselkoersfluctuaties. Het hoofdstuk sluit af met een zoektocht naar de redenen waarom de macro-economische component (het inflatieverschil) en de activa component (de nominale wisselkoers) van de reële wisselkoers zich zo verschillend gedragen in deze context.

Hoofdstuk 5 suggereerde dat reële wisselkoersen te gelijkmatig zijn vergeleken met fundamenten van de activa marketen (stochastische discontovoeten), terwijl hoofdstuk 6 suggereerde dat ze te volatiel zijn, vergelijken met de macroeconomische fundamenten (de relatieve consumptiegroei). Daarom bevat hoofdstuk 7 een gezamenlijke analyse van beide puzzels. Met behulp van dezelfde macro-economische dataset voor de 12 Eurozone landen die in hoofdstuk 6 gebruikt werd, onderzoekt het eerste deel van de empirische analyse in hoofdstuk 7 of de (volledige) verwijdering van de nominale wisselkoers resulteert in een reële wisselkoers volatiliteit die overeenkomt met de volatiliteit van de onderliggende consumptie processen. Voortbouwend op de discussie in dit eerste deel, presenteert deel twee van dit hoofdstuk een analyse van de 'dubbele scheiding' van de reële wisselkoers: aan de ene kant relatief ten opzichte van de activa markten, en aan de andere kant relatief ten opzichte van de macro-economische (consumptie) fundamenten. Daartoe wordt de macro-economische component (het inflatieverschil) afzonderlijk van de activa component van de reële wisselkoers (de nominale wisselkoers) geanalyseerd.

De empirische analyse in hoofdstuk 7 documenteert een brede 'scheiding', en een 'dichotomie' die bestaat tussen de macro-economische en de activa component van de reële wisselkoers. Ten eerst, suggereert het empirisch bewijs een duidelijke scheiding: het inflatieverschil blijkt te gelijkmatig te zijn, en onregelmatig gecorreleerd met de activa markt rendementen, terwijl de nominale wisselkoers te volatiel blijkt te zijn, en tevens onregelmatig gecorreleerd met relatieve consumptie. Ten tweede, lijkt iedere component op de juiste wijze samen te hangen met zijn fundamenten: het inflatieverschil hangt positief samen met de relatieve consumptiegroei (net zoals in hoofdstuk 6), terwijl de nominale wisselkoers positief samenhangt met het verschil in activa rendementen (hetgeen de risicovolle activa equivalent van de UIP conditie in hoofdstuk 3 ondersteunt).

Samenvattend, concentreerde dit proefschrift zich op het geven van empirisch bewijs rond de centrale vraag die aan het begin werd gesteld - hoeveel internationale risicodeling vindt er daadwerkelijk plaats. Een aantal alternatieve methodes, voortkomend uit verschillende theoretische raamwerken, suggereerden verschillende resultaten. Het meest macro-economische bewijs (in hoofdstukken 2 en 6) suggereerde dat er niet veel internationale risicodeling plaatsvindt. Daarentegen suggereerde de SDF methode (in hoofdstukken 4 en 5) dat er sprake is van bijna perfect internationale risicodeling. Nochtans komen alle methodes tot tenminste één gezamenlijk conclusie, los de vele onderlinge verschillen en discrepanties - de (reële) wisselkoers speelt een cruciale (essentiële) rol in alle maatstaven voor internationale risicodeling. In de meeste gevallen, lijkt de (reële) wisselkoers de 'stuwende kracht' te zijn achter de resultaten, die de gezamenlijke bewegingen van de onderliggende (consumptie of activa rendementen) processen goeddeels in de schaduw stelt. Daarom suggereert dit proefschrift, in het licht van de dichotomie rond het wisselkoersargument zoals beschreven in hoofdstuk 7, dat er wellicht een gemeenschappelijke noemer gevonden kan worden door een aanpak waarbij de hypotheses aan beide zijden van de dichotomie volledig separaat getest worden.

Dit proefschrift sluit af met een discussie over de meest relevante en toepasselijke maatstaven waarvan internationale financiële instellingen zoals het IMF gebruik zouden moeten maken om de voordelen van het verder openstellen van de globale kapitaalmarkt te evalueren, en ten einde landen van beleidsadvies dienaangaande te voorzien. Rekening houdend met de voornaamste bevindingen van de empirische hoofdstukken, concludeert dit proefschrift dat hiertoe verschillende maatstaven voor risicodeling zouden moeten worden gebruikt: beleidsadviezen zouden wellicht meer op macro-economisch bewijs gestoeld moeten zijn voor vaste/weinig flexibele nominale wisselkoersregimes, en waarschijnlijk meer op activa markten gebaseerde maatstaven voor flexibele wisselkoersregimes.

Резиме

Меѓународно споделување на ризикот е една од најважните потенцијални придобивки од процесот на меѓународна финансиска интеграција, кој се забрза во последните две децении. Сé додека различните национални економии не се совршено поврзани, постојат можности за споделување на ризикот преку меѓугранична трговија со средства и добра. Со собирање на ризикот во поголем, глобален пазар на капитал, жителите на овие економии можат да го диверзифицираат идиосинкратичниот, специфичен за земјата, дел од ризикот со кој се соочуваат, притоа стекнувајќи се со дополнителни добивки за благосостојбата коишто не се достапни во рамките на националните граници. Всушност, условот за совршено (оптимално) меѓународно споделување на ризикот сугерира целосна диверзификација на сите идиосинкратични (специфични за земјата) ризици, така што секоја земја се соочува само со (пропорционален дел од) агрегатниот (глобален), или системски ризик. На овој начин, агрегатните, недиверзифицирачки ризици би требало да имаат ист ефект насекаде низ светот, што во економскиот жаргон имплицира дека стапките на раст на маргиналната корисност би требало да бидат изедначени помеѓу сите земји.

Исто така, аргументите засновани врз потенцијалните придобивки од поголем и подлабок пазар на капитал беа често користени од страна на меѓународните финансиски институции со цел да охрабрат бројни земји во развој да ги отворат сопствените пазари на капитал и сé повеќе да се интегрираат со индустријализираниот свет. Неспорно, разумно е да се претпостави дека идното напредување на процесот на финансиска глобализација во голема мера ќе зависи од реалните придобивки за благосостојбата кои тој ги има испорачано во минатото. Токму затоа, прецизни процедури за квантифицирање на степенот на меѓународно споделување на ризикот се клучни за евалуацијата на процесот на меѓународната финансиска интеграција.

Сепак, алтернативните приоди кои може да се сретнат во академската литература сугерираат различни заклучоци во врска со степенот на споделување на ризикот помеѓу земјите, што всушност се случува. На пример, макроекономските методи, базирани врз серии на потрошувачка и производство, сугерираат ограничен степен на споделување на ризикот. Неколку опсервации го илустрираат нискиот степен на споделување на ризикот утврден во макроекономските студии: стапките на раст на (агрегатната) потрошувачка помеѓу ОЕЦД земјите се многу слабо поврзани, често се дури послабо поврзани отколку соодветните стапки на раст на производството; непостојаноста на потрошувачката е зголемена во однос на непостојаноста на производството за многу економии во подем, кои се вклучија во глобалниот пазар на капитал во текот на деведесеттите години; растот на релативната потрошувачка помеѓу две земји е типично аномално (негативно) поврзан со промените во реалниот девизен курс. Од друга страна, моделите за утврдување на цената на средствата, базирани на прекумерни приноси на пазарите на капитал, имплицираат многу висок степен на меѓународно споделување на ризикот. Главниот наод во овие модели е дека реалните девизни курсеви, кои служат како приближни вредности за разликата помеѓу стапките на раст на маргиналната корисност, се доста помалку непостојани од стохастичните есконтни фактори, кои пак се користат како приближни вредности за стапките на раст на маргиналната корисност во соодветните земји.

Поради тоа, земајќи го предвид очекуваното вклучување на дополнителни земји (во развој) во процесот на финансиска интеграција, важно е дополнително да се анализираат овие алтернативни методи во обид да се најде конечна рамка за квантифицирање на степенот на меѓународно споделување на ризикот.

Целта на оваа дисертација е да го истражи степенот до кој се случува меѓународно споделување на ризикот користејќи макроекономски и финансиски приоди, и да ги презентира наодите во компактна (синтетичка) рамка. Бидејќи е клучен за двата приода и природно ги поврзува, реалниот девизен курс презема улога на спроводник во оваа синтеза.

За да го одговори главното прашање - колкаво меѓународно споделување на ризикот се случува - оваа дисертација користи неколку различни, но сепак повремено меѓусебно испреплетени аспекти од макроекономската и финансиската литература. Суштинскиот дел на литературата поврзана со истражувањата преземени во оваа дисертација се состои од три аспекти: макроекономски модели од типот на реални бизнис циклуси, мерења базирани врз основа на портфолио теоријата и меѓународна рамка за утврдување на цената на средствата, базирана врз условите за непостоење на безризична заработка. Дополнително на овој суштински дел од литературата, истражувањата во некои поглавја се поврзани со еден или повеќе латерални аспекти: алтернативни канали за меѓународно споделување на ризикот (особено дознаки од работници од странство), тестови на условите за меѓународен паритет во приносите на пазарите на капитал и на каматните стапки, како и студии кои се однесуваат на одвојувањето на реалниот девизен курс од макроекономските фундаменти и (не)-неутралноста на режимите на номиналниот девизен курс.

Во поглавјето 2 се опишани алтернативните канали за меѓународно споделување на ризикот. Во првиот дел од емпириската анализа во ова поглавје се повторени некои наоди од литературата кои сугерираат дека меѓународното споделување на ризикот (во потрошувачката) е далеку од совршено. Користејќи податоци за најголем број на земји во развој и индустријализирани земји во периодот 1960-2000, панел анализата сугерира дека хипотезата на совршено споделување на ризикот во потрошувачката е отфрлена за секоја група на земји. Всушност, резултатите покажуваат дека стапките на раст на идиосинкратичната потрошувачка се силно поврзани со стапките на раст на идиосинкратичното производство, сугерирајќи дека само мал дел од севкупните макроекономски ризици се споделува помеѓу земјите. Освен тоа, а во согласност со најголемиот број емпириски студии, оваа врска е особено силна за групата на (помалку интегрирани) земји во развој, имплицирајќи притоа многу сериозно оддалечување од условот за совршено споделување на ризикот.

Вториот дел од емпириската анализа во поглавје 2 се фокусира на улогата на приливите на дознаки од работници од странство кон земјите во развој како алтернативен канал за меѓународно споделување на ризикот (во потрошувачката). Резултатите од панел проценките сугерираат дека земјите во развој со натпросечни приливи на дознаки од работници од странство (по глава на жител) исто така остваруваат релативно повисоки степени на меѓународно споделување на ризикот во потрошувачката (покажуваат релативно помали отстапувања од условот за совршено споделување на ризикот) во периодот 1990-2000. Сепак, и покрај тоа што оваа релација е точна за секоја подгрупа на земји во развој (повеќе финансиски интегрирани, помалку финансиски интегрирани и економии во транзиција), таа е статистички значителна само за групата на економии во транзиција. Признавајќи дека овој ефект би можело да се должи на неколку проблеми поврзани со податоците за оваа група земји (особено во текот на раните деведесетти години), анализата во поглавје 2 заклучува дека натпросечни степени на дознаки од работници од странство по глава на жител не водат до значително повисоко меѓународно споделување на ризикот во потрошувачката за земјите во развој.

Во поглавјето 3 е разработена можноста дека емпириската процедура која вообичаено се користи за тестирање на условот за меѓународен откриен паритет на каматните стапки (UIP)[1] би можела да биде пристрасна. Овој услов вели дека очекуваните приноси на две споредливи (безризични) средства деноминирани во различни валути треба да бидат еднакви, доколку овие средства се разликуваат само во однос на валутата на деноминација. Оттаму, овој услов се смета за еден од најчесто користените ценовно-базирани мерки за

[1] Uncovered interest rate parity (UIP) condition.

(интензитетот на) процесот на меѓународна финансиска интеграција. И покрај тоа што е теоретски добро поткрепен концепт, UIP е конзистентно отфрлуван во емпириските тестови. Ова поглавје почнува со демонстрација дека коефициентот од процедурата која вообичаено се користи за тестирање на UIP може да биде пристрасен доколку базичниот процес на генерирање на податоците незначително се разликува од теоретски очекуваниот процес.

Емпириските наоди во поглавје 3 сугерираат дека отфрлувањето на UIP условот би можело да биде ограничено на вообичаените тестирања кои ги регресираат промените во номиналниот девизен курс врз разликата помеѓу каматните стапки. На пример, спецификациите кои го регресираат приносот на странски депозити изразен во домашна валута врз домашната каматна стапка, вообичаено не успеваат да го отфрлат UIP условот. Всушност, овие спецификации обично создаваат коефициенти кои се блиску до теоретски очекуваната вредност од единица. Освен тоа, каматната стапка на земјата-сидро, која би можело да се смета за пропуштена варијабла во вообичаените тестови, влегува значително во основната спецификација. Според тоа, и двата наода даваат поддршка на хипотезата за пристрасност.

Исто така, преостанатиот дел од емпириската анализа во поглавје 3 ја истражува важноста на оваа пристрасност за проценката на коефициентите во вообичаените тестови на UIP. За оваа цел, таа презентира резултати од регресии за сите возможни (билатерални) парови на земји помеѓу 10 развиени, индустријализирани земји, и ги поврзува овие резултати со детерминантите на пристрасноста. И покрај тоа што специфичната структура на сетот со податоци не дозволува за директен тест на важноста на пристрасноста, наодите сугерираат дека главните детерминанти на пристрасноста се значително поврзани со проценките на коефициентите. Најпосле, поглавјето завршува со една забелешка за претпазливост: и покрај тоа што типот на естимациона пристрасност идентификуван во ова поглавје би можел да биде важен за вообичаените тестови на UIP, неговото земање во предвид не помага во рехабилитирањето на UIP условот, ниту во (целосното) решавање на UIP загатката.

Поглавјата 4 и 5 се фокусираат на мерење на меѓународното споделување на ризикот базирано врз приодот на стохастичен есконтен фактор (SDF).[2] Спротивно на бројните наоди од макроекономските студии базирани врз модели на меѓународни реални бизнис циклуси и од финансиските студии вкоренети во портфолио теоријата, овој приод сугерира висок степен на диверзификација на макроекономските ризици помеѓу земјите. Почнувајќи со основниот услов на (меѓународно) утврдување на цената на средствата, кој вели дека реалните девизни курсеви треба да бидат еднакви на разликата помеѓу движењата на стохастичните есконтни фактори за соодветните земји, билатералниот SDF приод покажува дека: прво, постои значителен макроекономски ризик кој би можело да се сподели помеѓу земјите бидејќи есконтните фактори се многу непостојани; и второ, многу голем дел од овој макроекономски ризик всушност се споделува помеѓу земјите бидејќи реалните девизни курсеви се многу помалку непостојани. Поглавјата 4 и 5 презентираат некои ограничувања и проширувања на овој приод.

Поглавје 4 презентира теоретска рамка за билатералниот SDF приод за меѓународно споделување на ризикот и реплицира некои од резултатите познати претходно во литературата. Освен тоа, ова поглавје идентификува две главни неконзистентности на билатералниот SDF приод. Прво, есконтните фактори во билатералната рамка не се единствено определени и решавачки зависат од другата земја вклучена во пресметките за споделување на ризикот. Второ, овие отстапувања, т.е. непрецизноста во мерењето на растот на маргиналната корисност, се најсериозни за земјите со релативно ниски прекумерни приноси на пазарите на капитал (ниски Sharpe ratios). Како обид да се земат предвид овие некохерентности, во поглавјето 4 е презентирано трилатерално проширување на SDF приодот. Меѓутоа, и покрај тоа што трилатералната рамка демонстрира дека финалните резултати за индексот на меѓународно споделување на ризикот се прилично робусни во однос на бројот на земји вклучени

[2]Stochastic discount factor (SDF) approach.

во нивната пресметка (индексите на споделување на ризикот се дури и повисоки во овој случај), таа не ги решава проблемите со внатрешната неконзистентност на билатералниот SDF модел.

Поглавјето 5 е продолжение на поглавјето 4 и укажува на тоа дека мерките за меѓународно споделување на ризикот базирани на SDF решавачки зависат од режимот и однесувањето на номиналниот девизен курс.

Всушност, емпириската анализа демонстрира дека SDF приодот механички или автоматски води кон совршено споделување на ризикот при епизоди на фиксни номинални девизни курсеви. Мерките базирани на SDF приодот имплицираат (буквално) совршено споделување на ризикот во две епизоди на фиксни номинални девизни курсеви: прво, помеѓу дванаесетте земји од Еврозоната после воведувањето на заедничката валута; и второ, помеѓу неколку земји во развој со фиксни номинални девизни курсеви во однос на американскиот долар и САД. Освен тоа, овој приод имплицира подобра интеграција во смисла на повисоко макроекономско споделување на ризикот помеѓу неколку земји во развој (дури и без целосно фиксен девизен курс) и САД отколку помеѓу развиени земји како Велика Британија или Јапонија и САД. Овој многу неинтуитивен резултат може да се смета за главно ограничување во користењето на SDF приодот за споредби во степенот на споделување на ризикот помеѓу различни земји. Земајќи ги предвид неуверливо високите резултати за степенот на споделување на ризикот добиени за овие епизоди, во поглавјето 5 се заклучува дека реалниот девизен курс би можело да биде навистина многу стабилен, но тоа не имплицира неопходно дека споделувањето на ризикот е совршено.

За разлика од претходните две поглавја базирани на пазарите на капитал, поглавје 6 се фокусира на општа макроекономска рамка за мерење на меѓународното споделување на ризикот. Тоа ја истражува важноста на флуктуациите во номиналниот девизен курс за аномалната врска помеѓу стапката на раст на релативната потрошувачка и промените во реалниот девизен курс. Оваа аномална релација се движи спротивно од условот за ефикасно меѓународно споделување на ризикот во потрошувачката извлечен од моделите на меѓународни реални бизнис циклуси, кои сугерираат дека потрошувачката треба да биде релативно повисока во земји кои доживуваат пад на релативното ценовно ниво. Според тоа, оваа аномална врска помеѓу потрошувачката и реалниот девизен курс сочинува една од главните загатки во меѓународната макроекономија, исто така позната и како Backus-Smith загатка. Главната цел на ова поглавје е да истражи до која мера оваа аномалија може да се припише на промените во номиналниот девизен курс. Во оваа смисла, елиминацијата на номиналните девизни курсеви помеѓу земјите од Еврозоната со воведувањето на еврото служи како своевиден природен тест за овој постоечки проблем.

Фокусирајќи се на дванаесетте земји од Еврозоната после воведувањето на заедничката валута (временски период 1999-2006), емпириската анализа покажува дека номиналниот девизен курс е навистина главната причина за оваа загатка - во отсуство на флуктуации во номиналниот девизен курс растот на релативната (билатерална) потрошувачка е позитивно поврзан со флуктуациите во реалниот девизен курс, што е во согласност со макроекономската теорија (модели на реални бизнис циклуси - RBC). Спротивставувајќи ги овие резултати со неколку алтернативни примери на флексибилни номинални девизни курсеви (дванаесетте земји од Еврозоната пред воведувањето на еврото (временски период 1986-1999), како и шест развиени, индустријализирани земји во текот на временскиот период 1986-2006), анализата демонстрира дека номиналниот девизен курс е главниот извор на Backus-Smith загатката. И покрај тоа што е сеуште позитивно поврзана со (негативната вредност од) разликата помеѓу стапките на инфлација во овие случаи, релативната потрошувачка е негативно поврзана со флуктуациите во номиналниот девизен курс. Следствено на тоа, поглавјето завршува со дополнително прашање во врска со причините за дивергентното однесување на макроекономската (разликата помеѓу стапките на инфлација) и на финансиската компонента (номиналниот девизен курс) на реалниот девизен курс.

Додека поглавјето 5 сугерираше дека реалниот девизен курс се чини премногу стабилен споредeн во однос на фундаментите од пазарите на капитал (стохастичните есконтни фактори), поглавјето 6 сугерираше дека тие се чинат премногу непостојани во однос на она што го имплицираат макроекономските фундаменти (стапката на раст на релативната потрошувачка). Токму затоа, во поглавјето 7 е извршена взаемна анализа на двете загатки. Користејќи го истиот макроекономски сет на податоци за дванаесетте земји од Еврозоната користен во поглавјето 6, првиот дел од емпириската анализа во поглавјето 7 истражува дали (целосната) елиминацијата на номиналниот девизен курс го доведува нивото на непостојаност на реалниот девизен курс во согласност со непостојаноста на базичните процеси на потрошувачка. Надоврзувајќи се на дискусијата во првиот дел, вториот дел од поглавје 7 понатаму го анализира двојното исклучување на реалниот девизен курс: од една страна во однос на пазарите на капитал, а од друга страна во однос на макроекономските фундаменти. За оваа цел, макроекономската компонента (разликата помеѓу стапките на инфлација) е анализирана одвоено од финансиската компонента на реалниот девизен курс (номиналниот девизен курс).

Емпириската анализа во поглавје 7 документира едно широко исклучување и една дихотомија, кои постојат помеѓу макроекономската и финансиската компонента на реалниот девизен курс. Прво, наодите сугерираат едно јасно исклучување: разликата во стапките на инфлација се чини премногу стабилна и аномално поврзана со приносите на пазарите на капитал, додека номиналниот девизен курс се чини премногу непостојан и исто така аномално поврзан со релативната потрошувачка. Второ, секоја компонента се чини дека е соодветно поврзана со фундаментите на својата страна: разликата помеѓу стапките на инфлација е позитивно поврзана со стапката на раст на релативната потрошувачка (слично на наодите во поглавје 6), додека номиналниот девизен курс е позитивно поврзан со разликата помеѓу приносите на пазарите на капитал (притоа давајќи поддршка на ризичниот еквивалент на UIP условот од поглавје 3).

Накратко, оваа дисертација се фокусира на приложување на емпириски наоди во врска со централното прашање покренато на почетокот - колку меѓународно споделување на ризикот се случува. Алтернативните приоди, вметнати во различни теоретски рамки, сугерираат дивергентни резултати. Најголемиот дел од макроекономските наоди (поглавја 2 и 6) сугерираат дека не се случува големо споделување на ризикот помеѓу земјите. Од друга страна, SDF приодот (поглавја 4 и 5) сугерира речиси совршено меѓународно споделување на ризикот. Сепак, независно од бројните разлики и разидувања помеѓу нив, сите приоди стигнуваат барем до еден заеднички заклучок - реалниот девизен курс игра стожерна (суштинска) улога во сите мерки за меѓународно споделување на ризикот. Всушност, во најголем број случаи, реалниот девизен курс се чини дека ги води резултатите за споделувањето на ризикот, притоа во голема мера засенувајќи ја важноста на взаемното движење помеѓу базичните процеси (на потрошувачка и приноси на пазарите на капитал). Токму затоа, имајќи го предвид аргументот за дихотомија на реалниот девизен курс елабориран во поглавје 7, оваа дисертација сугерира дека заедничка основа помеѓу различните приоди би можело да се бара во едно јасно и целосно одвојување на тестовите спроведени на секоја страна од линијата на оваа дихотомија.

Најпосле, дисертацијава завршува со дискусија за најрелевантните и најсоодветните мерки што меѓународните финансиски институции како ММФ би требало да ги следат во оценувањето на придобивките и во формулирањето на препораки за земјите коишто тежнеат кон (дополнително) отворање на нивните граници кон глобалниот пазар на капитал. Земајќи ги предвид главните наоди од емпириските поглавја, дисертацијава заклучува дека различни мерки (начини) за мерење на меѓународното споделување на ризикот би можело да се искористат за оваа цел: препораките би можеле повеќе да се засновуваат на макроекономските наоди во режими на фиксен или многу ригиден номинален девизен курс и веројатно повеќе на наодите од пазарите на капитал во случај на флуктуирачки номинален девизен курс.