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U.S.E. Working Paper Series nr: 21-04

School of Economics

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Utrecht University School of Economics

U.S.E. Research Institute
Working Paper Series 21-04
ISSN: 2666-8238

Market forces in healthcare insurance: the impact of healthcare reform on regulated competition revisited

Jacob A. Bikker[a,b,1]
Jack G.J. Bekooij[a]
[a]De Nederlandsche Bank, Amsterdam
[b]Utrecht School of Economics Utrecht University

January 2021

Abstract
This paper investigates the impact of market forces on competitive behaviour and efficiency in healthcare by investigating the Dutch healthcare insurance reform in 2006. This reform replaced the dual system of public and private insurance with a single compulsory health insurance scheme, in which insurance providers compete for customers in a free market. We measure competition directly from either shifts in market shares, or developments in profits. Using formal tests we find that in each approach a structural break occurs after the reform: competition is significantly higher after 2006 than before. Several robustness tests confirm this outcome. Nevertheless, we find that the health insurance sector is still less competitive than the banking, manufacturing and service industries, and even less competitive than life insurance.

Keywords: (regulated) competition, concentration, healthcare insurance, performance-conduct-structure model, Boone-indicator, scale economies.

JEL classification: G22, H51, L11-L13.

Comments welcomed to: j.a.bikker@uu.nl

1 The views expressed here are solely those of the authors and do not in any way represent the views of DNB

1. Introduction

In 2006, the Dutch healthcare insurance market underwent radical reform, with the aim of fostering efficiency and innovation by strengthening market forces (Enthoven and Van de Ven, 2007; Daley and Gubb, 2011). The dual system of public and private healthcare insurance coverage was transformed into one insurance system, with a mandate for individuals to purchase insurance, wherein healthcare suppliers compete for patients on a free market. Following the reform, the government no longer managed prices or volumes, but supervised quality and ensured a fair playing field. The government still sets outs the requirements for the basic health insurance package, and the Dutch Ministry of Health is responsible for calculating the annual accounting cost price of healthcare in the basic package. One pillar is solidarity: consumers pay a fixed insurance premium, but the system includes sliding-scale income-based subsidies for those with a low income as well as sliding-scale income-based taxation for those with a high income. A second pillar is accessibility: each consumer must be accepted by the insurer of its choice. For that reason, there is risk equalization for insurers. The impact of market forces on healthcare regularly leads to public debates. A key question is to what degree the new regime has indeed enforced the efficiency of, and the competition between, healthcare insurers.

To answer this question, we compare healthcare market conditions after 2006, which includes market forces, with the situation before 2006, without or with limited market power, and monitor the developments in market power over time. In this paper, market power refers to the role insurers play in lowering operational costs and profit margins, so they can lower healthcare insurance premiums. A key element of the reform is that insurers should use their influence to lower cost and raise quality of healthcare suppliers. We do not investigate the impact of the insurance reform on the cost and quality of healthcare.

This paper builds on an earlier study of Dutch healthcare reform by Bikker (2017). Compared to that study we have implemented four important improvements. First, we apply an improved version of the competition measurement model, which is essential to test changes in competition after the healthcare reform; the respective technical changes are explained later. Second,

we are now able to use disaggregated data, *i.e*. distinguishing basic insurance and supplementary insurance for more accurate and more extensive analyses. Third, we use pure health insurance data, as earlier data were polluted by observations of insurers with disability policies. And, finally, we have added five years to the post-reform period, almost doubling its length. Key results are much more convincing than the respective outcomes in Bikker (2017). This relates particularly to the annual estimates of competition which now reveal significantly more competition immediately after the start of the reform.

For our investigations we apply a competition measure, which we refer to as the performance-conduct-structure (PCS) model, introduced by Hay and Liu (1997) and Boone (2001, 2008). This model is based on the efficient-structure hypothesis, the idea that in a competitive environment, firms experience an increase in market share, if they pass on their efficiency gain (fully or in part) by lowering their output prices. Firms, in our application insurers, enjoy also higher profits, due to a larger market share and – if they retain part of their efficiency gains – a higher profit margin. In other words, efficiency is rewarded more highly amid heavier competition. For an overview of this PCS approach, see Bikker and Van Leuvensteijn (2015). This PCS measure has been employed in the past for the life insurance market (Bikker and Van Leuvensteijn, 2008; Bikker, 2016) and for health insurance (Bikker, 2017). The PCS indicator measures the extent to which existing efficiency across insurers is reflected in divergence of their performances. While there are alternative methods to measure competition, such as the traditional Lerner-index, the Panzar-Rosse model, concentration indices, and the price-cost margin model, most of them are hampered by data insufficiencies, theoretical flaws or empirical failings (Bikker and Spierdijk, 2017; Shaffer and Spierdijk, 2017).

This paper contributes to the literature by applying the PCS measure of competition, which has not been used frequently. A similar investigation of competition effects of a healthcare reform is rare or absent, except Bikker (2017) mentioned above. We use a unique, not publicly available data set for the Dutch health industry during 1995-2017, which captures the effects of the 2006 healthcare reform package. The results are interesting for other countries too, *e.g.* Germany and Switzerland, where similar reforms where undertaken (Busse

et al., 2017; Greß, Manouguian, and Wasem, 2007), the United States, particularly for policy makers who seek universal coverage (Rosenau and Lako, 2008) and all other countries where healthcare costs are rising and lowering healthcare prices are a continuous concern.

In addition, we estimate scale economies of health insurers. First, scale economies are frequently used as an indirect measure of competition. The underlying assumption is that strong competition incentivizes insurers to become more scale efficient, for example by forcing managers to reduce marginal costs in order to remain profitable (Raith, 2003; Hay and Liu, 1997). Persistence of unused scale efficiency would indicate the absence of strong competition (Kox and Van Leeuwen, 2011). Second, existing unused scale economies would be a key threshold for new entries on the health insurance market. Absence of the threat of new entries would reduce competition.

The remainder of this paper is organized as follows: Section 2 provides literature on competition and efficiency in healthcare insurance and section three gives, first, background information on the organization and development of the Dutch health insurance market and, second, an overview of empirical data of this market. Section 4 introduces the methodology of measuring competition as used in this paper. Section 5 provides annual estimates of health insurance competition based either on market share shifts or on net profit changes, while Section 6 tests whether competitive behaviour did indeed significantly increase after the 2006 reform, using various approaches. Section 7 presents scale economy estimates and test on a break in 2006. The final section provides concluding remarks.

2. Literature on competition and efficiency in healthcare insurance

Rosenau and Lako (2008) draw initial lessons for the United States from the new Dutch health insurance model. The first is that costs may not been controlled. In the earlier post-reform years, consumer premiums were increasing, and

insurance companies reported large losses on the basic policies[1]. Second, regulated competition is unlikely to make citizens happy and public satisfaction on market forces is not high. Third, consumers may not behave as economic models predict, remaining responsive to price incentives. Finally, policymakers should not underestimate the opposition from healthcare providers who define their profession as more than simply a job. These conclusions reflect voices that are heard regularly in later years, *e.g*. in politics and public debate.

ACM (2016) provides a survey on competition in the Dutch health insurance market after the 2006 reform. They follow a structure-conduct-performance (SCP) framework to examine the competitive qualities of the market and find that for healthcare insurers there are limited possibilities to distinguish themselves by product differentiation due to the extended regulation. They also observe little dynamic on the market, and absence of new market entrants. These three elements of limited competition were also found by Robinson (2004) for the healthcare insurance market in the US, which differs from the Dutch one, particularly in the sense that – on that market – both for profit and non-profit insurers operate. ACM (2016) mentions various reasons for the absence of new entrants: license application process, complexity of and high risks in the market, complexity and uncertainty in laws and regulations, and solvency requirements. In our view, the most important threshold for entries is scale economies, which we will address in this paper. Healthcare insurance is complicated and involves substantial fixed costs, which would be a burden for small, new insurers. Another cause that may limit competition is "The high number of different insurance policies can also lead to inertia among consumers, making them reluctant to switch", which may confuse consumers (so-called 'obfuscation'). The ACM report does not present a direct measure of competition. It also focuses only on basic healthcare insurance policies, due to data restrictions. There is greater freedom for product differentiation in terms of supplementary insurance, and we are able to investigate these supplementary insurance policies in this paper.

1 Although in the years 2006 and 2007 health carriers were accepting losses in order to build market share, more recently we have seen a steady increase in health insurance premiums (Swiss Re, 2011; Leu *et al.*, 2009).

Gaynor and Town (2011) remark in their book 'Competition in Health Care Markets' that until recently there have been very few studies on competition between health insurance firms (p. 83). While empirical research on the life and property and casualty insurance industry is more developed, we found very few studies on health insurers. In most countries, health insurance is included in life insurance. Cummins and Weiss (2014) list 74 studies treating various economic aspects of both non-life and life insurance, particularly their economic efficiency (including X-efficiency and scale economies), with roughly half of them studying the US market. Previous research into Dutch non-life insurance, which includes health insurance (and where health premiums written cover more than 50% of total premiums), revealed substantial scale economies averaging above 10% (Bikker and Gorter, 2011). Cummins and Weiss (2014) report substantial scale economies in studies focusing on the US and several European countries.

3. The Dutch health insurance industry

3.1. Structure

To obtain a better understanding of the Dutch health insurance industry, this section presents an overview of this sector and the main recent events shaping its development. Prior to January 2006, the Dutch healthcare system featured a complex structure of private and public insurance entities under the Compulsory Health Insurance Act (in Dutch: *Ziekenfondswet*), divided into three pillars. The first compartment is Basic health insurance: the mandatory National Health Service Institutions (NHSI) for all residents below the so-called NHSI income level. Covering 62% of the population, the NHSI was financed through income-dependent contributions paid by employees, employers and social security providers. Those not qualifying for the NHSI scheme could take out voluntary private health insurance at a flat-rate insurance premium. NHSI insured could expand the cover of their basic health insurance on the private health insurance market. A final compartment consisted of a public insurance scheme providing long-term care for the chronically ill, funded from social security premiums.

In 2006, the first two pillars were merged into a single private but mandatory scheme. Private insurers cover basic healthcare and compete in the market on price. All insured pay a flat rate for the basic package, while their employers pay

an income-dependent premium. The benefits are fully standardized and insurers are obliged to accept all applicants, regardless of their health profile (*i.e.* no 'cherry picking'). Insurers negotiate prices with healthcare providers and policyholders are free to change insurers every year (Schäfer *et al.*, 2010). There is risk equalisation whereby insurers with proportionally older or less healthy policyholders are compensated by insurers with relatively younger and healthy policyholders. The insured may supplement their basic package with extensions supplied by the market. The ultimate goal of the reform was 'to encourage health insurers to increase the efficiency of the healthcare provision by becoming prudent buyers of health services on behalf of their customers' (Van de Ven and Schut, 2009, p. 253).

There are now nine health insurance groups, at least so in the supplemental health insurance market, with 29 different brands. In 2006, there were 45 different basic health insurance packages and 137 supplemental health insurance policies on the market. Over the past year, these numbers have increased to 71 and 276, respectively. Each year 4 to 8% of policyholders switch their provider. Newly-introduced policies often had lower premiums than the average at the time.

3.2. Data used

This paper is based on data reported by healthcare insurers to their prudential supervisor, the Nederlandsche Bank. We use only data of monolines, insurers which provide one type of insurance product, here: healthcare insurance. Multiline insurers do not (and need not) report separate cost and profit data for their various types of products, so that their aggregated data are not useful for our investigations. After the 2006 reform, the share of monolines in the total healthcare premiums is above 90 percent. As we aim to model behaviour, it is not a problem that we do not include 10% of the market. Furthermore, we do not have data from the public insurance sector, consisting of mandatory NHSI, for the years before 2006.

Table 1. Developments in health insurance over 1995-2017

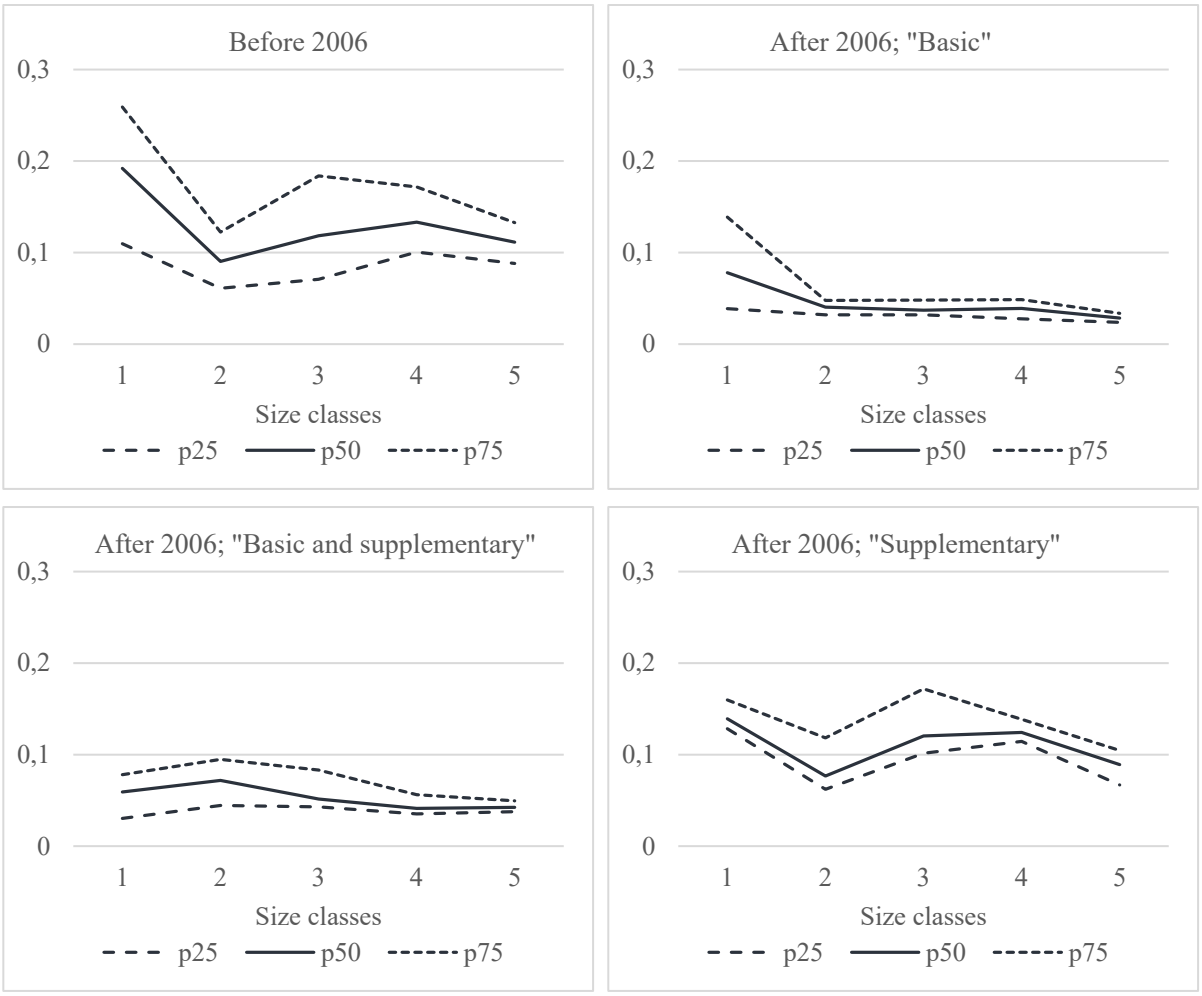
	1995	2000	2005	2006	2010	2015	2017
	Private market			*Total market*			
Gross premiums of insurers (mln euro*)	4.979	4.384	9.275	36.727	43.347	43.092	43.750
Idem, National Health Service Institutions (NHSI)	—	—	—	15.971	31.495	32.840	32.872
Total assets of insurers (mln euro*)	6.644	6.108	9.002	24.067	30.911	34.725	32.261
Idem, NHSI	—	—	—	10.130	21.026	23.288	22.159
Number of insurers, incl. NHSI from 2006	44	32	23	42	36	33	29
Concentration index HHI	669	732	1188	716	808	1045	1140
Profits/ Gross premiums, incl. NHSI from 2006	0.08	0.03	0.08	-0.01	0.02	0.02	-0.01
Operational costs/Gross premiums, incl. NHSI.	0.09	0.13	0.11	0.06	0.04	0.04	0.04

Source: DNB, own calculations;
Note: * Asterisks denote prices of 2017.

Table 1 shows developments in the healthcare insurance over 1995-2017. The left-hand side refers to the private market of 1995-2005, exclusive of the public sector, and the right-hand side to the entire market of 2006-2017. In 2006, the formerly public healthcare institutions (NHSI) enter the free competition market as new players. We see that in later years the number of insurers declines from 44 to 29, due to consolidations. The total relative operational costs of all insurers falls after 2006 from 11% to 6%. The profit margin fluctuates but, on average, also declines since 2006, and is even negative in a number of years. Based on the decline of both cost and profit margins from, on average, 15% to 6%, and on the total gross premiums of 2005, the cost savings of the private insurers alone amounts to more than 800 million euro annually.

Figure 1 and Table 2 present the operational costs as a percentage of gross premiums. First, we notice that the cost before 2006 is with, on average, 12% of gross premiums (the solid line in subfigure 1a) much higher than in later years where the percentage is 4 (the solid line in subfigure 1b). The lower cost reflects higher efficiency and may point to increased competition. Second, this figure reveals differences between basic and supplementary insurances. Insurers which provide both basic and supplementary insurances have cost margins of only 5% (subfigure 1c) against 12% for those with only supplementary insurance policies (subfigure 1d).

Figure 1. Operational costs as % of gross premiums for five size classes based on gross premiums, and for four types of insurance product



Note: p25, p50 and p74 refer to, respectively, the 25th, 50th and 75th percentile of the operational costs as % of gross premiums.

Third, Figure 1 shows that cost differences among insurers within a size class were much larger before 2006 than after that year: the percentile lines are close to each other after 2006. This may point to higher cost differences among insurers and, hence, to more inefficiency before the reform. Fourth, the figures suggest scale economies, as the costs of the class of smallest insurers are at least 40% higher than those of the class of large insurers (Table 2). This cost difference is even 174% for basic insurance after 2006.

Table 2. Average operational costs as a percentage of gross premium

	Before 2006			**After 2006**								
Type of insurance:	*Private insurers*			*Basic*			*Basic and supplementary*			*Supplementary*		
Percentiles	25%	50%	75%	25%	50%	75%	25%	50%	75%	25%	50%	75%
Size class 1 (small)	0.11	0.19	0.26	0.04	0.08	0.14	0.03	0.06	0.08	0.13	0.14	0.16
Size class 2	0.06	0.09	0.12	0.03	0.04	0.05	0.04	0.07	0.09	0.06	0.08	0.12
Size class 3	0.07	0.12	0.18	0.03	0.04	0.05	0.04	0.05	0.08	0.10	0.12	0.17
Size class 4	0.10	0.13	0.17	0.03	0.04	0.05	0.04	0.04	0.06	0.11	0.12	0.14
Size class 5 (large)	0.09	0.11	0.13	0.02	0.03	0.03	0.04	0.04	0.05	0.07	0.09	0.10
Average	0.09	0.12	0.17	0.03	0.04	0.05	0.04	0.05	0.08	0.08	0.12	0.14
Standard deviation		0.04			0.01			0.02			0.03	
Dispersal*		72			174			40			56	

Source: DNB, own calculations;

Notes: Size classes are based on gross premiums. Classes have equal numbers of insurers. * Dispersal denotes the difference in operational costs between smallest and largest size classes (respectively size class 1 and 5) as percentual mutation, based on original, not rounded figures.

4. Methodology of measuring competition

To measure competition among health insurers we estimate the so-called performance-conduct-structure (PCS) indicator of competition (Bikker and Van Leuvensteijn, 2014). This indicator is based on the concept that more efficient insurers on a fiercely competitive market will gain a larger market share, if they pass on at least part of their cost advantage onto their premiums:

$$\ln MS_{jt} = \alpha + \beta_t \ln MC_{jt} + \varepsilon_{jt} \qquad (1)$$

The dependent variable $ln\,MS_{jt}$ represents the market share of insurer *j* in year *t,* in logarithm form to correct for heteroscedasticity. The coefficient of the logarithm of marginal costs MC_{it} is β_t, an elasticity, which is the PCS indicator of year *t*. This indicator is expected to have a negative sign, as more efficient firms will obtain higher market shares. In absolute terms, low negative values are interpreted as weak competition, and *vice versa*, while 0 would mean: no competition at all. An alternative is to have a constant indicator value: $\beta_t = \beta$ for all *t*, or different (but constant) values before and after the 2006 reform. We may estimate efficiency as average cost (AC), which in practice works as well as model-based estimated marginal costs (Bikker, 2017, p. 70-71). As the market shares add to one each year, we substitute the restriction $MS_{pt} = 1 - \sum MS_{jt}$ for

each year (summing over 1, 2, …, p –1) into the model equation by dividing each observation by that of the p^{th} insurer:

$$\ln(MS_{jt}/MS_{pt}) = \alpha + \beta_t(\ln AC_{jt}/AC_{pt}) + \varepsilon_{jt} \quad (2)$$

Following Hay and Liu (1997) and Bikker and Van Leuvensteijn (2008), we estimate additional models with net profit as an alternative performance indicator:

$$\ln(P_{jt}/P_{pt}) = k + \lambda_t(\ln AC_{jt}/AC_{pt}) + \varepsilon_{jt} \quad (3)$$

Note that profit can be seen as market share times profit margin. The profit measure captures the idea that the industry rents are an inverse function of competition, reallocating profits to the most efficient firms and proving the selection effect of competition. Although it is acknowledged that other unobserved sector-specific factors may affect the PCS indicator, we can, within bounds, compare PCS indicators across industries, sectors, and over time in order to assess the level of competitive strength.

5. Annual estimates of competition in healthcare insurance: the market share model

5.1. Market shares

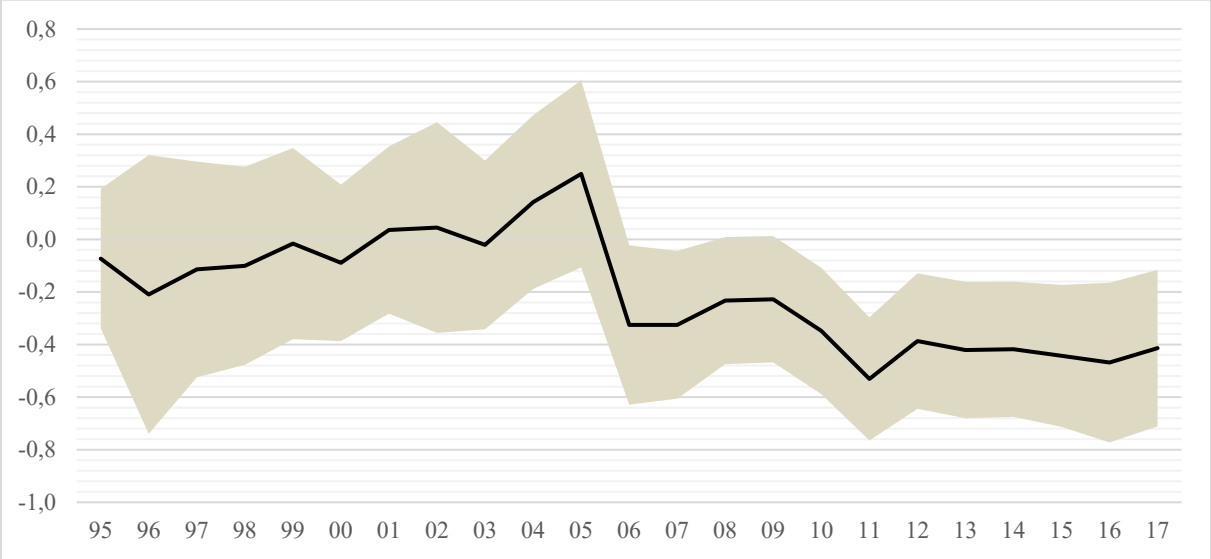
To capture the developments of competition in healthcare insurance over time, we estimate annual values of the PCS indicator over 1995-2017 using Equation (2) with a small adjustment:

$$\ln(MS_{jt}/MS_{pt}) = \alpha_1 + \alpha_2 + \beta_t(\ln AC_{jt}/AC_{pt}) + \varepsilon_{jt} \quad (4)$$

Note that a_1 applies for all t before 2006 and a_2 for all t since the beginning of 2006. In 2006, after the healthcare reform, the number of market participant increased as former public health institutions entered the private sector market. The model constant reflects the average market share, which fell after 2006. Hence Equation (4) has a pre-2006 constant, a_1, and a post 2006 one, a_2. The estimation results are presented in Figure 2 and Table 3. First, we notice the

strong downwards development in 2006, pointing to a jump to heavier competition, where lower operational costs of health insurers result more strongly in larger market shares. This effect becomes even stronger in later years. Second, we observe that competition is statistically significant, as the 95% confidence interval lies fully below the y-axis, except for 2008 and 2009 where that statement holds only for the 90 % confidence interval. Before 2006, competition has not been significant in any year. Hence, the health insurance reform has significantly increased competition between healthcare insurers.

Figure 2. Effect of operational costs on market shares: the PCS indicator (1995-2017)



Note: The black line presents the PCS indicator over time, the shaded area shows the 95 percent confidence intervals.

This graph differs from the similar figure 4 in Bikker (2017), where the sharp jump towards competition after the reform was absent, or even showed a contrary movement. This was due to a less sophisticated application of the PCS model. Bikker (2017) wrongly chose to have one constant over the entire period, but here we have instead two period dummies, one for the years before 2006 and the other for the years thereafter, in order to reflect the different level of market players, or the varying average market shares.

Table 3 presents the estimation results of the competition measure as shown in Figure 2. In addition, we provide the outcomes for the weighted regression, where each observation is weighted by the size of the health insurer in terms of

gross premium, so that each euro paid in premiums counts equally. The results are fairly similar, although the level of significance is slightly lower.

Table 3. Annual estimates of competition in healthcare insurance (1995-2017)

	Market shares		**Profit**	
	Unweighted	*Weighted*	*Unweighted*	*Weighted*
1995	-0.07	-0.05	-0.69**	-0.95***
1996	-0.21	-0.08	-1.06**	-0.73*
1997	-0.11	-0.07	-0.35	-0.63*
1998	-0.10	-0.07	-0.47	-0.77**
1999	-0.02	-0.01	-0.48	-0.69
2000	-0.09	-0.14	-0.31	-0.46
2001	0.04	-0.16	0.27	0.04
2002	0.04	-0.25	-0.30	0.14
2003	-0.02	-0.25	-0.75	-0.54
2004	0.14	0.01	0.14	0.32
2005	0.25	0.13	0.23	0.44
2006	-0.33**	-0.42**	-0.63	-0.09
2007	-0.32**	-0.41**	-0.37	-0.38
2008	-0.23*	-0.29*	-0.85***	-0.72**
2009	-0.23*	-0.28	-1.17**	-0.53*
2010	-0.35***	-0.42**	-0.63*	-0.41
2011	-0.53***	-0.46**	-0.84**	-0.52
2012	-0.39***	-0.42**	-1.10***	-0.82***
2013	-0.42***	-0.45**	-1.00***	-0.67*
2014	-0.42***	-0.44**	-0.98**	-0.70**
2015	-0.44***	-0.49**	-0.56	-0.30
2016	-0.47***	-0.49**	-0.94**	-0.28
2017	-0.41***	-0.35*	-1.21***	-0.73**

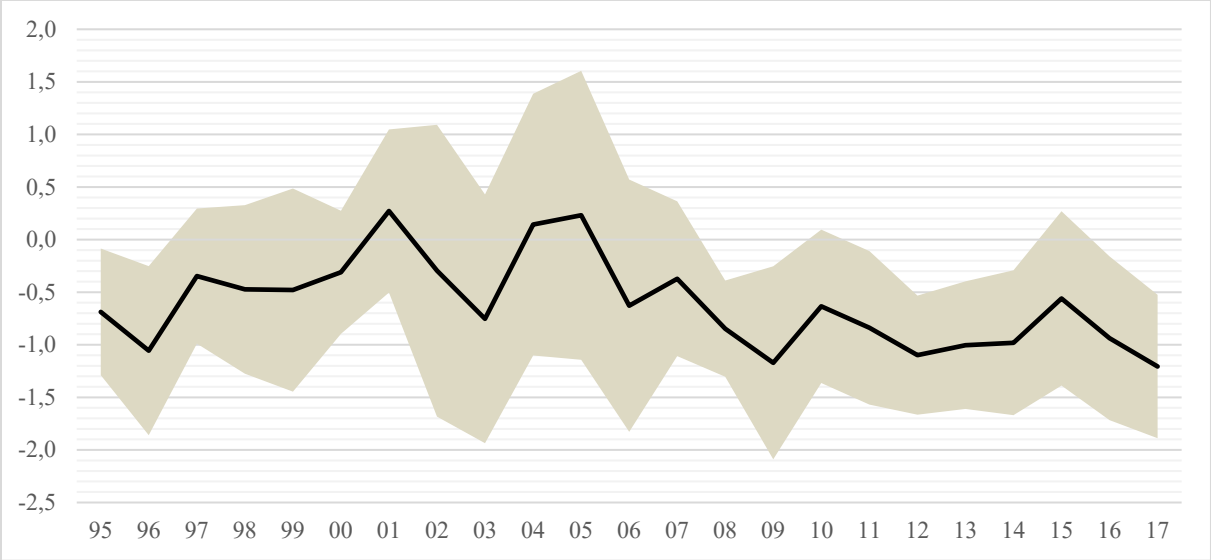
Notes: Since 2006 the data set includes former NHSI. Asterisks *, ** and *** denote significant difference from zero at, respectively, the 90%, 95% and 99% confidence level.

5.2. Profits

We have repeated the above annual analysis, but now for profits, based on Equation (3; where a has been replaced by a_1 and a_2) also used frequently in the literature (Boone, 2008). The impact of operational costs on profits consists of two combined effects: where insurers pass on cost advantages in the form of lower prices, they obtain a larger market and hence generate more profits (profit margin times sales) and to the degree that cost advantages were retained, the profit margin increases. Figure 3 shows how the advantages of lower operational costs on profits of health insurers over time becomes stronger, pointing to more competition. In the reform year 2006, the Boone indicator increases less prominently than in the market share analysis. But again it appears that before 2006 the competition is not statistically significant (at least since 1997), whereas

this is always the case after 2008 (see right-hand panel in Table 3). Four remarks must be made in this respect. First, in 2006 and 2007, healthcare insurers were competing heavily on the new market by offering insurance policies below cost price. Thereafter, insurance premiums increased to the benefit of insurers' profits. Second, profits are much more volatile than market shares, so that a clear demarcation line is less visable. Third, separate regressions over the subperiod before and after 2006, presented below, show a clear structural break around 2006 (see below). Finaly, Table 3's right-hand column shows weighted regression estimates of Equation (3). The outcomes are similar to those of the unweighted regression, with slightly lower significance levels.

Figure 3: The impact of operational costs on profits of healthcare insurers by year



Note: The black line represents the PCS indicator over time, the shaded area gives the 95% confidence interval.

6. Tests on a possible structural break in reform year 2006

6.1. Separate regressions

The upper panel of Table 4 presents the estimations using the following equation:

$$\ln(MS_{jt}/MS_{pt}) = \alpha_1 + \alpha_2 + \beta_1(\ln AC_{jt}/AC_{pt}) + \beta_2(\ln AC_{jt}/AC_{pt}) + \mu_j + \varepsilon_{jt} \quad (5)$$

Note that compared to Equation (4) $\beta_t = \beta_1$ for all t before 2006 and $\beta_t = \beta_2$ for all t since the beginning of 2006. In fact, this are two separate regressions, one for the period before the health insurance reform, and the other for the years thereafter, similar to a Chow structural-break test. We apply the fixed-effects (FE) estimation approach. The insurer-specific fixed-effects μ_j pick up persistent relevant differences across insurers, which may affect market shares or net profits and, hence, avoid omitted variable bias.[2] Furthermore, we apply Newey-West to estimate heteroskedastic and autocorrelation corrected (HAC) standard errors (and *t*-values). The uppermost panel of Table 4 presents the results for market shares, using unweighted regressions. The key outcomes of annual estimates of competition in the section above are repeated: the impact of average cost on market shares is not significant before 2006 (pointing to absence of competition) and significant at the 99% confidence level after that year for the all insurers sample (indicating a certain degree of competition). Apart from unweighted regressions (with equal weights for each insurer) we also apply tests based on weighted regressions, where each euro payed in premiums counts equally (2nd panel of Table 4). Both estimation variants have comparable test results.

After 2006 we have data for insurers which provide either (i) only basic health insurance, (ii) both basic and supplementary, or (iii) only supplementary. Basic insurers operate in a competitive market: the coefficient of average costs is statistical significant (3rd column of Table 4). For basic & supplementary insurers as well as for the supplementary insurance case, we do not find significant values for the PCS indicator. One explanation is that competition may fall short because insurers may offer supplementary packages of their own composition, allowing for product differentiation. Another explanation may be that these samples are too small to show significant results.

2 A random-effect model was rejected by the Wu-Hausman test, in favour of the fixed-effect model.

Table 4. FE estimates for the healthcare market share model (1995-2017)

	Pre 2006	**Post 2006**			
		All insurers	*Basic*	*Basic & supplem-entary*	*Supplem-entary*
	(1)	(2)	(3)	(4)	(5)
	Market shares model, unweighted				
Average costs, ln	-0.06	-0.27***	-0.27**	-0.09	-0.47
# of observations	380	436	212	119	105
# of insurers	57	50	29	10	11
	Weighted				
Average costs, ln	0.08	-0.28**	-0.17**	-0.21	-0.82
	Profit model, unweighted				
Average costs, ln	-0.39*	-1.45***	-0.94***	-1.70*	-2.48***
# of observations	284	297	140	78	79
# of insurers	56	50	29	10	11
	Weighted				
Average costs, ln	-0.55**	-1.67***	-1.23***	-2.19***	-2.73**

Notes: Asterisks *, ** and *** denote significant difference from zero at, respectively, the 90%, 95% and 99% confidence level. Numbers of observations and numbers of insurers are identical for the weighted and unweighted cases.

The lower panel of Table 4, repeats the estimations for profits instead of market shares. We observe that the PCS indicator is not significant for before 2006 and significant at the 99% confidence level after that year for the all insurers sample, similar as in the market share case. For the weighted regression estimates we see significance at the 95% confidence level for the pre 2006 sample. What is crucial here for our test of a structural break in 2006, is a jump in competitiveness of the 'all insurers' case after 2006 compared to before 2006, which is reflected in the much larger value of the PCS indicator value, in absolute terms. Most results in this table are in line with Bikker (2017).[3]

For this profit model we find the highest PCS indicator values (in absolute terms) for the supplementary insurers, where consumers are free to choose a supplementary package, while the mixed insurers, providing both types of insurance, take a middle position. This is remarkable, as insurers have more product differentiation opportunities on the supplementary market. The small

3 An exception is the pre-2006 market share case in Table 4 (comparable to Table A.4 in Bikker (2017). This data collection was polluted by observations of insurers with a less exclusive healthcare nature, such as disability insurance.

numbers of observations of these groups call for caution with respect to conclusions.

6.2. F-tests on structural breaks

Table 5 presents an extended series of F-tests on the difference in competition before and after the reform year 2006. The tests are based on Newey-West estimates of heteroskedastic and autocorrelation corrected (HAC) standard errors. The first row shows tests based on the 'separate regressions' or Chow structural break model of Equation (5): α and β different for both sub-periods – corresponding with a structural break in 2006 – versus α and β identical across the entire sample period, reflecting absence of a break. The second row presents results for the 'model with two period dummies' from Equation (4): one regression with α and β different for both sub-periods – a structural break – versus α different for both sub-periods and only one β for the entire sample period – no structural break. This second model focus only on the competition indicator β.

Table 5. F-tests on higher competition between health insurers after 2006 compared to before 2006 (1995-2017)

	Market shares				**Profits**			
	All insurers		*Excl. NHSI*		*All insurers*		*Excl. NHSI*	
	Un-weighted	Weighted	Un-weighted	Weighted	Un-weigh-ted	Weighted	Un-weighted	Weighted
Separate regressions or Chow structural break	10.15*** (0.0000)	24.05*** (0.0000)	7.15*** (0.0009)	19.62*** (0.0000)	7.46*** (0.0006)	2.85* (0.0586)	6.58*** (0.0015)	4.45*** (0.0021)
Number of observations	[816]		[610]		[581]		[427]	
Model with two period dummies	10.09*** (0.0015)	21.97*** (0.0000)	4.20** (0.0408)	6.31** (0.0122)	0.52 (0.4692)	0.15 (0.6960)	1.44 (0.2302)	3.81* (0.0517)

Notes: P-values in parenthesis. Number of observations in square brackets. *, ** and *** denote significant difference from zero at, respectively, the 90%, 95% and 99% confidence level.

The first two columns of the 'market shares' panel in Table 5 relates to the 'all healthcare insurers' sample. Before 2006 this group concerns the private sector insurers only, as we have no data of the NHSIs. After the 2006 reform the sample increases with the former NHSI institutions which now enter the private market. Market behaviour of the insurers may have changed after the 2006

reform but – at the same time – the composition of insurers has changed. We have no reason to expect that the public sector mandatory NHSIs were more competitive than the private sector insurers, on the contrary in fact. Nevertheless, to disentangle the effects of these two (possible) changes, we re-estimate all regressions exclusive of the former NHSIs, so that the F-test on the structural break is than based on exactly the same sample before and after 2006. The third and fourth columns give the outcomes for the private health insurers only, thus exclusive of the NHSIs. Six out of eight cases have F-tests which are significant on the 99% confidence level in proving the existence of a structural break, while two out of eight cases have F-tests which are significant on the 95% confidence level.

The right-hand panel of Table 5 presents F-tests of the profit model. The profit-based tests confirm a structural break in 2006 (first row), but the tests focussing on the competition indicator β do not prove a boost in competition after 2006. The profit-based PCS model may be inferior to the market-share based model for theoretical reasons. As said above, profit is the product of market shares and profit margins. Lower marginal (or average) costs can be used fully or in part to gain a larger market share (which would point to the existence of competition), but an insurer may also hold the cost advantage and add it to its profit. This would raise the profit margin, but this behaviour is not what is expected under heavy competition. Though the profit-based PCS model is used often in the literature, our conclusion is that profit is less suitable as dependent variable in the PCS model, compared to market shares.

6.3. The dynamic PCS model

Following Hay and Liu (1997), we introduce as another robustness test a one-year lag of the market share variable to capture lagged adaptation: a permanent decline in marginal costs may have a gradual upward effect on market shares. Therefore, a positive coefficient is expected on the lagged term:

$$\ln(MS_{jt}/MS_{pt}) = \alpha + \beta\,(\ln\, AC_{jt}/AC_{pt}) + \gamma\,(\ln MS_{j,t-1}/MS_{p,t-1}) + \mu_j + \varepsilon_{jt} \qquad (6)$$

We exclude for our sub-period regressions observations of 2006, because for that year, we do not have lagged market shares for either new healthcare insurance

entrants (such as the formerly NHSI) or merged healthcare insurers. Note that in a lagged model the long-term effect of marginal costs on market shares is $\beta/(1-\gamma)$. The upper panel of Table 6 shows that, for the market share model, gradual adaption is indeed a fact: the lagged market share is highly significant and the annual adaptation is only 40% (following from $(1-\gamma)=0.4$). [4] Lagged adaption is absent for profits (2nd panel of the table). The long-term results for $\beta/(1-\gamma)$ confirm the outcomes of the non-dynamic model. First, for market shares and unweighted profits, we find significant long-term competition after 2006, and absence of that before that year. Second, for weighted profits, we observe a huge – significant – jump in the competition measure $\beta/(1-\gamma)$ after 2006.

4 Gradual adaptation of market share indicates the likelihood of autocorrelation in the errors and is, hence, an argument to apply Newey-West (for autocorrelation corrected) estimates of the standard errors, as we did throughout this paper.

Table 6. FE estimates for the dynamic healthcare insurance model (1995-2017, excl. 2006)

		Unweighted		**Weighted**	
		Pre 2006	*Post 2006*	*Pre 2006*	*Post 2006*
		Market share model			
Average costs, ln	*β*	0.05	-0.27*	-0.03	-0.25
Market share, lag, ln	*γ*	0.59***	0.64***	0.66***	0.64***
Long-term AC	*β/(1-γ)*	-0.13	-0.76**	-0.10	-0.69*
# of observations		321	381	321	381
# of insurers		54	49	54	49
		Profit model			
Average costs, ln	*β*	-0.36	-0.82**	-0.56**	-0.93**
Profits, lag, ln	*γ*	-0.04	0.15	-0.02	0.12
Long-term AC	*β/(1-γ)*	-0.35	-0.97**	-0.55**	-1.06**
# of observations		193	207	193	207
# of insurers		50	42	50	42

Notes: Asterisks *, ** and *** denote significant difference from zero at, respectively, the 90%, 95% and 99% confidence level. The long-term competition effect is calculated as *β*/(1-*γ*). After 2006, the NHSI are included.

6.4. Comparison of competition across financial institutions

An absolute benchmark for the effect of marginal costs on market shares or profits is absent. In order to judge the intensity of competition, we need to compare our results with similar estimates from other industries. In their study of the banking sector, Van Leuvensteijn *et al.* (2011) use the market share PCS model and find that the *β* indicator averages -2.5 in the long-run, which is much higher – in absolate terms – compared to our healthcare insurance values ranging from -0.27 to -0.76 (Tables 4 and 6).[5] Bikker (2016) investigates the life insurance industry using a dynamic model similar to ours, which facilitates comparison. For the FE estimate of the long-term effect, he finds a value of -0.92, slightly higher in absolute terms than what we observe for healthcare insurance. Creusen *et al.* (2006) estimate the PCS model based on profits for the Dutch manufacturing and service industries and found elasticities between average variable costs and profits of around -5.7 and -2.5 respectively, above – in absolute terms – our FE health profit estimate ranging from -0.94 to -2.48 (Table 4). Hence, we conclude that the health insurance sector is less competitive than the banking, manufacturing and service industries, and even less competitive than life insurance.

5 Van Leuvensteijn *et al.* (2011) estimate a model without the lagged 'market share' so that *β* is their long-term PCS indicator.

7. Scale efficiency of healthcare insurers

As an additional analysis we estimate unused scale economies. Scale economies and other inefficiencies act as an indirect measure of competition, as under heavy competition we do not expect (large) scale economies or high inefficiencies. Furthermore, substantial existing scale economies would hamper new entrants and, hence, the impact of their possible threat on existing insurance firms. We use a translog operational cost function of output, input prices and other relevant variables as a second-order Taylor expansion around the mean, in natural logarithms (see Bikker, 2017). We present the complete estimation results in Table A.1 of the appendix, where we also provide the regression model and comment on the outcomes. Table 7 applies a constant returns to scale (CRS) test on the scale economies estimates of that operational cost model. Scale economies for the average sized insurer was close to zero for the years until 2006, but after 2006 scale economies were much higher by 14%: expansion of production goes with costs that are 14% lower than those of the current production. This is in line with the estimates of Bikker (2017). Where health insurance in the past implied, generally speaking, checking claims and compensating these expenses (which are variable costs), nowadays there are many costs that are related to negotiations with suppliers of medical care and the development of a more complex, strategic policy required for healthcare insurers, involving activities which primarily have a fixed nature – not or hardly varying with size. These higher costs needs to be allocated to large numbers of insured clients. This requires a large scale, and that is indeed what we observe in the falling number of insurers since 2006. It also demonstrates that there are large impediments for potential new entrants.

Table 7. Scale economies and CRS tests on the translog operational cost models (1995-2017)

	Since 2006			**Before 2006**
	Basic	*Basic and supplementary*	*Supplementary*	
Incl. NHSI	0.143 ***	0.111 **	0.018 **	0.0004 ***
(Number of observations)	(189)	(109)	(96)	(380)
Excl. NHSI	0.124 ***	0.115	0.085	
(Number of observations)	(83)	(65)	(85)	

Note: Asterisk *, ** and *** denote significant difference from zero at, respectively, the 90%, 95% and 99% confidence level.

Table 7 presents scale economies for various data sets ('the fracture numbers'), as well as their CRS tests (the 'asterisk'). Before 2006, we found 0.04% scale economies for the average healthcare insurer, though – remarkably – still significantly different from zero. Since 2006, we observe 14.3% scale economies for 'basic' health insurance, for the full sample, including former NHSI. Note that the fixed costs of negotiations and policy development argument, mentioned above, typically regards basic insurance policies. Scale economies for 'supplementary' is very low at 1.8% (though significant). Apparently, supplementary insurance is more straightforward, comparable to the healthcare insurance before 2006. For 'basic and supplementary' we find a weighted average of these numbers (11.1%).

8. Conclusions

Competition among health insurers has increased significantly since the 2006 reform of health insurance. We find this result particularly for the market share structure-conduct-performance model. These results are robust for various models, various samples (including and excluding NHSIs), weighted and unweighted regressions, and dynamic and non-dynamic model specifications. One key goal of the healthcare insurance reform has been achieved: more competition. Furthermore, the operational costs of health insurers have fallen since the reform, particularly with respect to the basic healthcare insurance. Most likely, the enforced influence of market power has improved cost efficiency of health insurers, among others, by consolidation. Although we observe statistical

significant improvements, the health insurance sector is still less competitive than the banking, manufacturing and service industries, and even less competitive than life insurance.

The change in the role of the health insurers has strongly increased their unused economies of scale since 2006. One the one hand, this stresses the need for a consolidated market with large players, but on the other, it points to impediments for new market entrants. From cost considerations, further consolidation could be desirable, but taking the need for competition into account, it is necessary that an ample number of players remain active on the health insurance market.

The quality of healthcare is ultimately the major key goal of the healthcare insurance reform. However, that topic is outside the scope of our economic investigations.

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APPENDIX. SCALE ECONOMIES ESTIMATES FOR HEALTH CARE INSURANCE

Economies of scale exist when the average cost of production is negatively related to a firm's output. For a service based industry, such as insurance, one has to rely on value measures. Cummins and Weiss (2014, pp. 26-33) report an extensive debate in the literature about the most appropriate measures of output. Keeping this discussion in mind, we follow Bikker and Gorter (2011) and consider (gross) premiums as a possible measure of the insurance service of covering normal risks or expected losses. Additionally, we include total assets as second output measure, representing financial services. Total assets generate investment income while those assets also act as a buffer for lagged claims, unexpected losses, and future healthcare spending due to aging of insured client populations . We estimate scale economies using a translog cost function, see Bikker (2017):

$$\ln OC_{jt} = \alpha + \sum_i \beta_{Yi} \ln Y_{ijt} + \frac{1}{2} \sum_i \sum_k \gamma_{Yik} (\ln Y_{ijt} - \ln Y_{i\bullet\bullet}) (\ln Y_{kjt} - \ln Y_{k\bullet\bullet}) + \sum_k \delta_k X_{kjt} + u_{jt} \quad (7)$$

OC_{jt} is the total operational cost of health insurer j in year t (t = 1, 2, …, T), defined as the sum of management cost (or administrative cost) and acquisition cost (that is, marketing and sales cost), and Y_{ijt} is output volume of type i (i = 1, 2; premiums and total assets). Operational costs and output terms are expressed in logarithms, which reduces heteroscedasticity and generates elasticities as coefficients. The model contains squares and cross-terms of output components in order to pick up any non-linearity in the cost elasticities – and hence economies of scale – across different size categories. All output types in the non-linear terms are expressed in deviation of their averages (in logarithms), calculated over all insurer-year combinations, *cf.* the Taylor series expansion. The average for output type i is denoted as $\ln Y_{i\bullet\bullet}$, with dots for the sub-indices over time and across insurance firms. The four other explanatory variable (X_k) are (i) the composition of operation costs in terms of management and acquisition cost ('distribution ratio'), (ii) the share of stock-based insurers, (iii) the Herfindahl-Hirschman Index (HHI), which reflect the structure of the market in terms of concentration, sometime interpreted as measure of competition, and (iv) a time trend. We express overall ray scale economies (SE) for insurer j in year t as:

$$SE_{jt} = 1 - \sum_{i=1}^{N} (\partial \ln OC_{jt} / \partial \ln Y_{ijt}) = 1 - \sum_{i=1}^{N} (\beta_{Yi} + \gamma_{Yii} (\ln Y_{ijt} - \ln Y_{i\bullet\bullet})) \quad (8)$$

The SE for the *average* health insurer is equal to (1 – $\sum_i \beta_{Yi}$), the sum of linear output elasticities. In that case, the squared terms disappear due to the fact that the log outputs are presented in deviation from their geometric averages.[6]

[6] This is the first simplification which is due to the functional form in Equation (7) of the non-linear output terms, that is, in deviation from the respective (geometric) mean. The second is that the cross-output terms in Equations (8) did disappear entirely, after taking first derivatives.

Table A.1. Scale economy estimates of health insurance based on premiums

	After 2006						Before 2006
	Basic		*Supplementary*		*Basic and supplementary*		
	Incl. NHSI	Excl. NHSI	Incl. NHSI	Excl. NHSI	Incl. NHSI	Excl. NHSI	Excl. NHSI
Gross premiums (in logs)	0.770***	1.048***	1.151***	1.204***	0.592***	0.756***	0.616***
Ditto, squared [a]	0.839	0.532	0.027	-0.377*	-0.276	0.570	0.012
Total assets (in logs)	0.087	-0.172	-0.169*	-0.289**	0.297*	0.129	0.384***
Ditto, squared [a]	1.116	0.860	-0.093	-0.175	0.083	0.539	-0.111
Cross term GP & TA [a]	-1.894	-1.195	0.154	0.510*	0.163	-1.107	0.104
Distribution ratio	0.039	0.349	0.119	-0.215	1.302***	1.170***	0.622***
Stock insurers	0.065	0.109	0.302**	0.308**	-0.316***	–	0.212***
HHI/100	-0.001	-0.004	-0.005	0.001	0.002*	-0.010	-0.001***
Time	-0.046***	-0.040**	-0.048***	-0.052***	-0.068***	-0.071***	0.011
Scale economies (SE)	0.143	0.124	0.018	0.085	0.111	0.115	0.000
Number of observations	189	83	96	85	109	65	380
R^2, adjusted (in %)	90.4	77.5	91.2	94.7	96.6	95.9	84.1
F test on CRS [b]	11.32***	8.88***	2.79**	1.78	2.49**	1.19	6.55***

[a] Squared values of gross premiums and total assets and their cross term are in deviation from their respective average values; [b] Critical value of the CRS test statistic (with four restrictions) at 5% and 1% significance ranges from, respectively 2.38 to 2.39 and 3.34 to 3.36, depending on the degrees of freedom.
Notes: We dropped the constant in this table. The indices *, ** and *** denote significant difference from zero at, respectively, the 90%, 95% and 99% confidence levels. Total costs, outputs and the wage index are expressed in 2017 prices.

Estimates of Equation (7) are presented in Table A.1. Gross premium is the dominant size variable and highly significant different from zero for all samples. The second size measure, total assets, was also highly significant before 2006, but fluctuate after that year. The two variable are, of course, correlated, but this does not impair our SE estimation approach as the coefficients of the first five variables together determine SE results. The SE outcomes for average firms, based on Equation (8), are explained in Section 7.

The F-tests in Table 7 on the null hypothesis of 'constant returns to scale' are based on the following four restrictions, applied to Equation (7). The sum of the two coefficients of gross premiums and total assets is one ($\beta_{Y1} + \beta_{Y2} = 1$) and the coefficients of both squared terms and the cross term are all zero ($\gamma_{Y11} = \gamma_{Y22} = \gamma_{Y12} = 0$). Rejection of the null hypothesis means that non-zero SE exist. As a robustness test, we dropped the distribution ratio. This did not affect the main results presented above.