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CHILD CARE PRICES AND MATERNAL EMPLOYMENT: A META-ANALYSIS

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Abstract. The literature estimates for labor force participation elasticity with regard to child care prices are extensive and varying. While some estimates imply substantial gains from child care subsidies, others find insignificant effects. To determine the causes of the variance, this paper reviews and analyzes the elasticity sizes using estimates from 36 peer-reviewed articles and working papers in the literature. We start by reviewing the theoretical and empirical aspects related to participation elasticity with regard to child care costs, paying special attention to sample characteristics, methodological aspects, and macro level factors. We conclude by providing a meta-regression using control variables based on our review of the literature to explain some of the differences between the estimates. As research builds on and improves the methods and assumptions in prior works, elasticity estimates have become smaller over time. This decline might also be partially explained by changes in labor market characteristics. In countries with high rates of part-time work and very high or very low rates of female labor force participation, we find elasticity rates to be smaller.

Keywords. Child care prices; Child care subsidies; Labor force participation; Maternal employment

1. Introduction

Child care subsidies are an important element of modern welfare states. It is an attractive policy option, given that it might reallocate resources to young parents and stimulate maternal employment.[1] To test the benefits of child care subsidies, since the late 1980s a sizeable body of literature has appeared investigating the impact of child care prices on maternal employment. This paper reviews the elasticity estimates from this literature, discussing its findings and attempting to explain why elasticity estimates differ so significantly across studies.

A puzzle arises when considering the empirical findings over time and across countries. While earlier studies, such as Blau and Robins (1988), find a large elasticity of maternal employment with regard to child care prices in the United States, there is a large variation in results when more recent studies from the United States and also Europe are taken into account. These findings challenge the hypotheses relating high maternal employment with child care subsidies. As an example, Bettendorf *et al.* (2015) find that an increase in the subsidy rate of child care had a modest impact on aggregate labor supply in the Netherlands. A striking example of variation over time can also be seen within US studies. The

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elasticity for single mothers calculated by Kimmel (1995) is close to −0.35 while Herbst (2010) estimates an elasticity of −0.05 for the same group.

A number of explanations can be offered for the variation in estimated elasticity sizes. A first set of explanations refers to study-specific factors such as methodology employed or sample characteristics; a second set takes into account macro-level factors such as the overall female labor market participation rate or the flexibility of the labor market measured by part-time work incidence. Aggregate indicators and institutional factors are inevitably ignored in micro-level studies since they usually show no variation. A meta-analysis of the results found in the literature can scrutinize both groups of explanations for variation in elasticity estimates, since we can compare results from different labor markets and samples. We study estimates from 11 countries and spanning 30 years between 1988 and 2010, the period in which a large number of child care price elasticity of employment were estimated using mainly structural methods.

The rest of the chapter is organized as follows. Section 2 provides a brief overview of the basic labor supply models with child care. Section 3 discusses the literature findings and describes the methodologies used, the differences in sample characteristics and the macro-level trends in elasticity estimates over time and across countries. Section 4 presents meta-regressions, testing the various explanations offered in Sections 2 and 3 for the variation in elasticity sizes. Section 5 concludes.

2. Theoretical Framework

The importance of child care prices for female labor supply was recognized by Heckman (1974) referred to as the "pioneering study in the field" by Blau and Robins (1988). The most commonly found theoretical basis for analyzing the effect of child care prices on labor supply in the literature is based on the simple static labor supply model with modifications to include child care prices and choices. A number of variants of the static model have been proposed with varying degrees of flexibility by including child care quality, unpaid care, and household decision making (Blau and Robins, 1988; Connelly, 1992; Ribar, 1995; Powell, 1997; Tekin, 2007). The starting point of labor supply models involving child care is a utility function consisting of consumption C and leisure l. The maximization problem has two constraints; time $l + h = 1$ and budget $C = y + (w - p)h$, where y is nonlabor income, h is hours of work, and p is price of formal child care. In this basic budget constraint, the price of child care is treated like a tax on wages with each hour of work requiring the purchase of an hour of child care. The implications of changes in the price of child care are thus similar to a change in taxation, with higher prices making it less likely for the effective wage rate $w - p$ to exceed the reservation wage.

Two common modifications to the basic model are unpaid care and quality of child care. The addition of unpaid care to the model relaxes the assumption that hours of formal child care use must be equal to working hours and provides an explanation for the observation of working women who do not use formal care. Instead, as in Blau and Currie (2006), the sum of unpaid care hours u and formal child care hours h^f may be assumed to be equal to working hours, $u + h^f = h$. Working hours will then also include any time spent in transit to and from work and child care. If the costs of unpaid care are assumed to be equal to the caregiver's market wage, the shadow price of unpaid care could be added directly into the budget constraint. A simpler approach is to add the leisure l_c of the unpaid caregiver directly into the utility function. Subject to the time constraint of the informal caregiver $u + l_c = 1$, there is a disutility associated with the use of unpaid care. If child care quality is equal across care types, the price of child care can be related to unpaid care with the condition $\frac{U_{l_c}}{U_u} = p$. This equality implies that the marginal cost of an extra hour of formal care is equal to that of informal care. The price of formal care can then be used as an indicator of the shadow price of informal care.

A second modification is based on relaxing the assumption that the quality of child care is homogenous among maternal care and other care types, by defining and including in the utility function a term for quality given by $q = (1 - h, C)$. It is assumed that care quality is improved by maternal care and consumption goods. Following Ribar (1995), the resulting maximization problem from these two modifications is given by

$$\max_{h,h_f} U(C, 1-h, 1-l, q) \tag{1}$$

subject to the consumption constraint $C \leq y + hw - h^f p$ and time constraints $l + h = 1$, $l_c + u = 1$ and $h^f + u = h$. Manipulating the first-order condition of this problem with respect to h, the reservation wage can be defined as

$$w_r \geq p + \frac{U_l + U_q Q_l}{U_c + U_q Q_c} \tag{2}$$

Equation (2) is similar to the result that Ribar (1995) derives with slight differences in the way unpaid care and child care prices are defined. Evaluated at the corner of the budget constraint where $h = 0$, the comparative statistics are self-evident. Child care prices, marginal utility of leisure for the mother, and the informal caregiver as well as higher quality maternal care raise the reservation wage and lower the likelihood of employment. High marginal utility from consumption and a large marginal contribution to child quality from consumption goods lower the reservation wage.

3. A Review of the Empirical Literature

We surveyed the literature in three steps. The sample was constructed between December 2010 and November 2011. First, the Google Scholar and EconLIT search engines were used, searching for the key phrases "labor supply child care elasticity" and "labor force participation and child care prices" as well as variations of these key phrases that replace "labor supply" with "employment" and child care with "child care" throughout the construction period. The second source of literature was the reference lists of the articles found in the initial search. Third, the literature review of Blau and Currie (2006) and the literature review of Wrohlich (2006) were extensively used, the former for mostly studies from the United States and Canada, and the latter for studies from Europe. Forty-four estimates are included in Table 1 for analysis from around 36 English language articles published between 1988 and 2010.[2] The studies estimating only the hours elasticity such as Heckman (1974) and Averett *et al.* (1997) were excluded since the comparison with participation elasticity is not possible in terms of the elasticity sizes.

Table 1 shows the calculated elasticity, sample size, year of the data used, country, and the data source for each article. The data sources used in the literature tend to be surveys that include information about child care expenditures and the samples sizes are usually below 10,000. Studies based on natural experiments are an exception because administrative and census data without micro-level information on child care costs can be used for natural experiment type studies and the resulting sample sizes are considerably larger (Lundin *et al.*, 2008; Cascio, 2009). There are a few cases such as Blau and Robins (1991), Del Boca *et al.* (2004), Wetzels (2005), and Van Gameren and Ooms (2009), where there is no elasticity estimate reported because the coefficient on price is close to 0 and statistically insignificant. In these cases, we set the elasticity to 0. Furthermore, two subsample estimates where elasticity sizes are reported based on statistically insignificant effects are also set to 0: the singles subsample elasticity reported by Kimmel (1998) and the married subsample elasticity reported by Cascio (2009).

It is common in meta-analyses to weight estimates using a precision factor, usually the standard error. The basic premise for the weighting is that more precise estimates will be closer to the true effect. The second use for the precision factor is to test for publication bias; the potential bias resulting from researchers only publishing statistically significant effects in the expected direction. If estimates' precision is significantly related to the effect size in one direction or the other, there can be said to be evidence of publication bias. While using the estimated elasticity rather than a regression coefficient as the effect size makes a comparison between studies convenient, none of the studies in our sample report a standard error

Table 1. Literature Summary.

Authors/publishing year	Elasticity	Sample size	Year	Country	Data source	Sample
Doiron and Kalb (2005)	−0.02	5305	1997	Australia	CCS SIHC	Married
Doiron and Kalb (2005)	−0.075	1116	1997	Australia	CCS SIHC	Single
Gong *et al.* (2010)	−0.287	4184	2006	Australia	HILDA	All
Powell (1997)[a]	−0.38	9201	1988	Canada	CNCCS	Married
Baker *et al.* (2005)[b]	−0.236	33708	1998	Canada	NLSCY	Married
Cleveland *et al.* (1996)	−0.388	3976	1988	Canada	CNCCS	Married
Wrohlich (2004)[c]	−0.025	1345	2002	Germany	SOEP	All
Wrohlich (2006)	−0.02	3213	2002	Germany	SOEP	Married
Del Boca *et al.* (2004)	0	1259	1998	Italy	Bank of Italy	Married
Van Gameren and Ooms (2009)	0	737	2004	Netherlands	SCP	Married
Wetzels (2005)	0	865	1995	Netherlands	AVO	Married
Kornstad and Thoresen (2007)	−0.12	2400	1998	Norway	IDS	Married
Lokshin and Fong (2006)	−0.46	1505	1999	Romania	RCCES	All
Lokshin (2004)	−0.12	2169	1995	Russia	RLMS	All
Gustafsson and Stafford (1992)[d]	−0.063	166	1984	Sweden	HUS data	Married
Lundin *et al.* (2008)	0	683471	2002	Sweden	Statistics SE	Married
Jenkins and Symons (2001)	−0.09	1235	1989	UK	LPS	Single
Viitanen (2005)	−0.138	5068	2000	UK	FRS	Married
Anderson and Levine (1999)	−0.511	12458	1992	USA	SIPP	All
Anderson and Levine (1999)	−0.463	9045	1992	USA	SIPP	Married
Baum (2002)	−0.185	2801	1991	USA	NLSY	All
Baum (2002)	−0.584	691	1991	USA	NLSY	Low income
Blau and Hagy (1998)	−0.2	2426	1990	USA	NCCS	All
Blau and Robins (1988)	−0.38	6170	1980	USA	EOPP	Married
Blau and Robins (1991)	0	10670	1984	USA	NLSY	All
Cascio (2009)	−0.505	68827	1950−1990	USA	Census	Single
Cascio (2009)	0	225028	1950−1990	USA	Census	Married
Connelly (1992)	−0.2	2781	1984	USA	SIPP	All
Connelly and Kimmel (2003b)	−0.37	1523	1992/3	USA	SIPP	All
Connelly and Kimmel (2003a)	−0.446	1523	1992/3	USA	SIPP	Married
Connelly and Kimmel (2003a)	−0.984	1523	1992/3	USA	SIPP	Single
Han and Waldfogel (2001)	−0.35	30931	1992	USA	SIPP	Married
Han and Waldfogel (2001)	−0.615	10187	1992	USA	SIPP	Single
Herbst (2010)	−0.05	74042	1999	USA	CPS/SIPP	All
Hotz and Kilburn (1991)	−0.01	2032	1986	USA	NLSY	All
Kimmel (1995)	−0.346	348.5	1987	USA	SIPP	Low income
Kimmel (1998)	0	697	1987	USA	SIPP	Single
Kimmel (1998)	−0.923	2350	1987	USA	SIPP	Married
Ribar (1992)	−0.74	3738	1984	USA	SIPP	All
Ribar (1995)	−0.088	3769	1984	USA	SIPP	All
Rusev (2006)	−0.25	3060	1997	USA	SIPP	All
Tekin (2007)	−0.121	4029	1997	USA	NSAF	Low income

(*Continued*)

Table 1. *Continued.*

Authors/publishing year	Elasticity	Sample size	Year	Country	Data source	Sample
Michalopoulos and Robins (2000)	−0.156	13026	1989	USA/Canada	CNCCS	Married
Michalopoulos and Robins (2002)	−0.259	1762	1989	USA/Canada	CNCCS	Single

[a]The author has a later article that focuses on the effect of prices on the use of different modes of care using the same dataset, Powell (2002). No employment elasticity is provided in that case although own price elasticities for child care modes are shown. Own calculations show that employment elasticity is above −1 if all care types are taken into account but around −0.16 if only center care effects are calculated. The latter value is also calculated by Doiron and Kalb (2005).
[b]This paper was later published as Baker *et al.* (2008), but the final version does not seem to include the elasticity calculation. The coefficients for employment seem to be the same in both versions. We use the elasticity estimate reported in the working paper.
[c]The labor supply model is estimated for a much larger sample but not all households have young children. The price selection equation is based on 1345 observations for children under six.
[d]The elasticity estimated is for working for more than 30 hours a week. The results in the meta-regression are largely similar if this study is dropped from the sample.
Bold entries in the author/publishing year column indicate working papers.

for the elasticity to be used as a precision factor. The standard error for the coefficient on price could be converted into a standard error for some studies, but this is not feasible for a large portion of the sample, given the characteristics of the studies in our sample. First, there is not a single price coefficient in papers using multinomial logit models. There is instead a coefficient for each category. There is no obvious way to transform the standard errors of these coefficients into a standard error for the price (nearly 30% of our estimates). Second, natural experiment type studies (9% of our estimates) do not have a standard error for a price or expenditure type variable. Third, several studies (e.g., Wrohlich, 2006) calculate the impact of a change in child care subsidies through its effect on consumption and do not have a price variable. Finally, several studies, namely Connelly and Kimmel (2003a), Michalopoulos and Robins (2000), and Michalopoulos and Robins (2002), report only significance levels and no standard errors at all. As a result, we follow the suggestion of Stanley (2005) and use the sample size as our precision variable and to test for publication bias, which is similar to the method of Card *et al.* (2010).

Table 1 shows the elasticity estimates across the literature. Bold entries indicate working papers while the rest are published papers. The child care price elasticity of labor supply ranges from nearly 1 in Kimmel (1998) and Connelly and Kimmel (2003a) to close to 0 in Lundin *et al.* (2008) and Wrohlich (2006). Below, we discuss the possible causes for the variation.

3.1 *Sample Characteristics and Methodological Choices*

The graph presented in Figure 1 shows the variation across elasticity estimates by sample size. The studies based on larger data sets appear to have smaller elasticity estimates and the resulting asymmetric distribution might indicate publication bias. However, there are also significant differences in the sample make-ups and methodologies of studies, which might explain the differences in elasticity sizes by sample size. We show some of these sample characteristics in Table 1. The findings with regard to marital status are somewhat ambiguous. Among the studies that estimate elasticity values for both groups Doiron and Kalb (2005) and Han and Waldfogel (2001) find single women to be more responsive, while Kimmel (1995, 1998) finds married women to be more responsive. Michalopoulos and Robins (2000, 2002) have investigated both groups and have also found a smaller elasticity for married women. For samples with lower income, the estimated elasticity by Kimmel (1995) for single women is larger than her later study

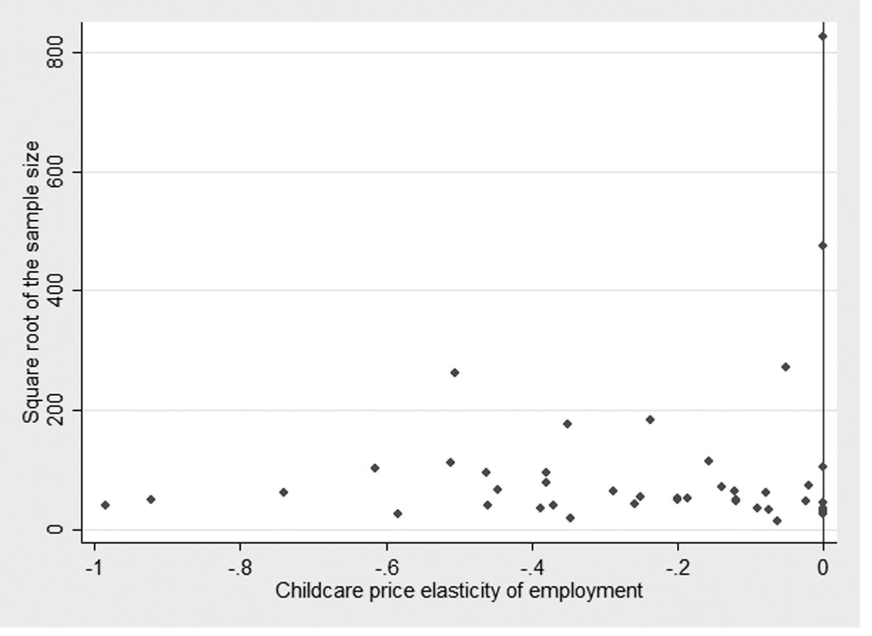

Figure 1. Sample Size and Elasticity Estimates.

for a general sample (Kimmel, 1998). Baum (2002) also finds a relatively large elasticity above 0.5 for a sample with low income. The only exception is Tekin (2007), who finds a relatively low elasticity for a low income sample, but this is more likely due to the methodology used in that study, which seems to produce generally smaller estimates.

Methodological choices can in fact play a large role in explaining the differences between the estimates. Many studies included in this analysis use a structural model to predict wages and child care prices, which are in turn used to estimate the employment elasticity using a limited dependent variable model such as probit (see, for example, Connelly, 1992; Ribar, 1992; Kimmel, 1998; Anderson and Levine, 1999). Beyond endogeneity issues in directly estimating the labor supply equation, wages and often formal care prices may not be available for nonworking mothers. Heckman (1979)'s sample selection specification is often used for the wage equation to avoid selection biases in OLS regressions. The selection term in the second stage is not always statistically significant in price and wage equations as seen in Wetzels (2005), Viitanen (2005), and Cleveland *et al.* (1996).

Gong *et al.* (2010) raise the issue that many of the earlier studies such as Connelly (1992) and Powell (1997) have to rely on user provided data to arrive at a price of child care. The standard approach is to divide reported child care costs with working or child care hours to arrive at a price of child care. The definition used to calculate the price of care can lead to different elasticity estimates. Although Kimmel (1998)'s original estimations use price per hours worked, the estimated elasticity becomes much closer to the value found in Ribar (1992) when the price variable is shifted to price per hour of child care utilized. Some studies such as Blau and Hagy (1998), avoid user-provided data by using additional data from child care providers.

Besides the structural models with two outcomes, employment or nonemployment, several scholars such as Michalopoulos and Robins (2000, 2002), Lokshin (2004), and Tekin (2007) use multinomial choice models. The basic setup involves defining the combinations of full-time or part-time work and unpaid or paid care as multiple discrete outcomes. The number of these outcomes depends on the categorical detail that the study reaches into, with Tekin (2007) using seven and Michalopoulos and Robin (2000, 2002) using 12 categories. The econometric *caveat* in these models is the possible correlation between the error terms in the estimation of the different indirect utility equations. The standard multinomial logit model imposes the assumption of independence of the irrelevant alternatives, meaning that the error terms should be uncorrelated. However, this assumption may be too strict for female labor supply and child care applications. Women with strong tastes for leisure may be less likely to work and have lower wages. Similarly, women who work full-time may prefer formal care over informal care. Mixed logit models can allow for correlations across the error terms to overcome this problem. Blau and Hagy (1998) and Tekin (2007) use discrete error structures similar to those found in duration models based on Heckman and Singer (1984). The estimated elasticities of multiple choice models are uniformly in the lower end of the estimates, a remark also made by Tekin (2007).

Apart from structural modeling, the second broad identification methodology in the literature is natural experiments. These studies exploit changes in child care policies to identify the effects. Although they do not identify as many parameters as structural models, reduced-form natural experiments may be more useful in identifying the effects of price changes in public or universal child care systems where there is little variation in price. Studies by Lundin *et al.* (2008), Baker *et al.* (2005), and Cascio (2009) fit into this category. A few studies also use subsidies explicitly within a structural model to estimate price effects, but the overall modeling strategy in these cases does not differ from other structural studies based on child care expenditure data (Rusev, 2006; Tekin, 2007; Herbst, 2010).

3.2 *Country and Time Trends in Elasticity Estimates*

With regard to the macro-level factors, the elasticity estimates in the literature show some visible patterns over time and across regions. The difference between estimates from Europe and the United States is quite large. Taking the mean of the two subsamples shows a mean of −0.15 for the European and Canadian studies and −0.35 for the US studies. Figure 2 shows the decline in elasticity sizes over time. The question is whether this decline is due to a change in women's responsiveness to child care prices. Limiting the comparison to the United States, Tekin (2007) and Herbst (2010), who use samples from 1997 and 1999, respectively, find much smaller elasticities than estimates based on samples from the 80s or early 90s. However, it is difficult to conclude directly that the smaller elasticity findings are merely due to changes in the population's responsiveness to child care prices, given the fact that there are also methodological differences between the studies. Furthermore, labor market conditions and institutions that are taken as given in micro studies may have changed. In particular, we test the relationship between child care price elasticity and three aggregate indicators: FLFP, part-time work and the level of inequality.

Borck (2014) provides a model where countries' institutional characteristics can lead to different equilibria in terms of FLFP and child care. In one equilibrium there is low participation and little demand for child care provision, while in a second equilibrium the reverse is true. Using European data, Lippe and Siegers (1994) find that women in very traditional networks are unlikely to respond to changes in wages. As child care subsidies provide a similar incentive, their effects can be influenced by social norms and the definition of gender roles. Going back to equation (2), this may be because parents in such countries view maternal care as having a higher quality than nonmaternal child care. Countries with low FLFP may then also have low child care price elasticity. However, there can also be a limit to how far child care prices can affect participation. Lundin *et al.* (2008) find insignificant effects from child care prices on maternal employment in Sweden and argue that further reductions in child care prices will have diminishing effects both because the prices are already low and because maternal employment is already high. On the aggregate level, higher labor force participation decreases the elasticity size simply because the

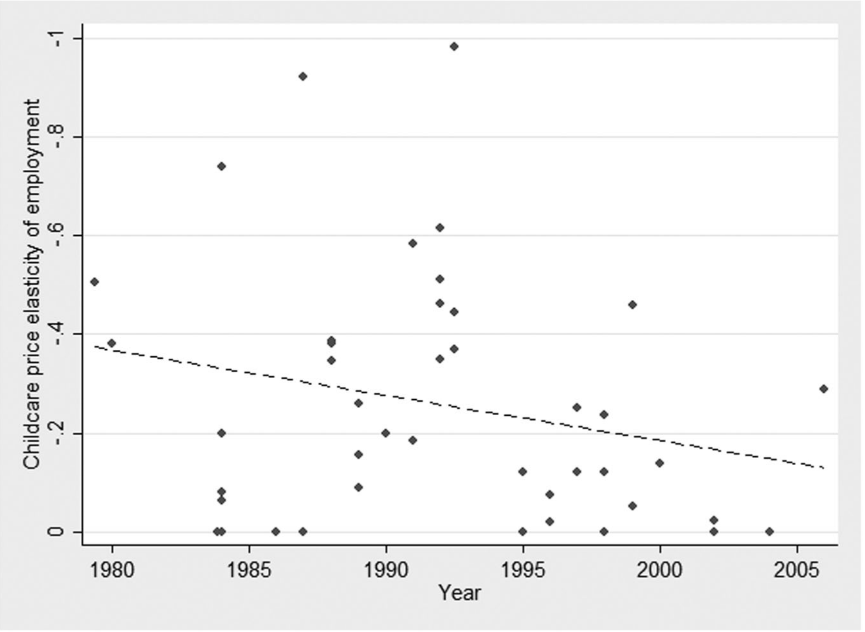

Figure 2. Elasticity Estimates Over Time.

denominator in the elasticity calculation becomes larger. High participation figures can also signal factors influencing the individual employment decision. Countries with high FLFP may already have cheap and readily available formal child care or can make use of alternative informal care arrangements, both of which will diminish the effects of further reductions in child care prices. Furthermore, if other structural factors, such as the wage structure or working hour flexibility, contribute to high female employment, the effect of the child care prices on employment may be limited at high levels of participation.

Another plausible source of variation in elasticity estimates is the prevalence of part-time work. Working part-time lowers the demand for formal, full-time child care, while at the same time allowing women to provide informal child care. These two effects of part-time work are expected to increase both the demand for and the supply of informal child care which can be substituted for formal child care. The substitution of informal care for formal care has been offered as an explanation as to why previous literature reviews find that the price elasticity of demand for formal child care is much larger than the labor supply elasticity with regard to the price of child care (Blau and Currie, 2006). If potential caregivers have more leisure due to part-time work, given that the benefit of providing an hour of care needs to equal the cost of losing an hour of leisure, the shadow price of informal care is expected to be low. Since the price of formal care equals this shadow price at the margin, a higher proportion of informal care is to be expected in settings with high part-time incidence. European estimates are supportive of this line of reasoning. Using an expansion in subsidies in the Netherlands where part-time work is prevalent, Bettendorf *et al.* (2015) have found a very large increase in formal care use but smaller effects on employment. The result suggests that the lower prices led to a large switch from informal to formal care.

A final macro-level factor which we might expect to influence the size of the elasticity parameter is the level of income inequality. While no functional form to the utility maximization problem was assigned in equation (1), the common assumption of concavity and hence diminishing marginal returns to consumption and leisure may have implications for the elasticity size. For example, if only low income mothers respond strongly to child care prices, the level of income inequality can determine the overall size of the elasticity.

4. Meta-Regressions

In this section, we use multivariate regressions to further investigate the impact that various methodological or macro-level factors may have on the elasticity estimates. Meta-regressions are quite varied in the literature, ranging from simple OLS models to random effects models where the effects are weighted by the inverse of variances (Nelson and Kennedy, 2009). In their review of meta-analyses in environmental economics literature, Nelson and Kennedy (2009) suggest that weighting and correction for heteroskedasticity are crucial for meta-regressions. Stanley and Doucouliagos (2013) suggest that weighted least squares (WLS) (with inverse of the standard error as weights) are preferable to more widely used random effects estimators. Since we do not have standard error information for all estimates, we estimate WLS regressions using the square root of the sample size as the weight variable.[3] We also regressed the elasticity sizes on the inverse of the square root of the sample size in a bivariate WLS specification and found a strong negative relationship. This seems to indicate that there might indeed be publication bias that we should control for. We therefore add the inverse of the square root of the sample size as a control variable in most specifications in addition to a dummy indicating Social Sciences Citation Index (SSCI) status. The estimated regression specification is given as

$$\beta_i = X'\gamma + Z'\phi + \varepsilon_i \tag{3}$$

In equation (3), β_i is the elasticity estimate of study i, X are controls for sample and methodological characteristics including the inverse of the square root of the sample size, and Z represents the macro-level factors which include a dummy for studies from the United States[4] and the year of the sample the study uses. ε_i is the error term. Since the sample is made up of both working papers and journal articles, we include a control for papers that have appeared in SSCI journals. To take basic sample differences into account, we add a control for estimates from a low income sample and the median value of the child age range used in the study.[5] For the methodological differences, controls for natural experiments and multinomial models are added. Once we control for natural experiments and discrete choice models, the base category is mostly made up of probit and ordered probit estimates. We considered adding a variable for the number of controls used but constructing a harmonized variable for the number of control variables is difficult. Most controls are quite standard such as education levels and age, and different definitions for these variables (such as continuous or category variables) can translate into large differences in the number of controls. Furthermore, methodology and sample characteristics often determine the inclusion of some control variables such as taste parameters and year fixed effects. It is not surprising to find that natural experiments have fewer control variables. Finally, two German studies, Wrohlich (2004) and Wrohlich (2006) use the same data source, similar methodologies, and find similar results. In this case, we simply took the average of the two estimates, which are quite similar in size, and took the mean of their sample sizes to weight the resulting average estimate. Summary statistics for the variables can be found in Table 2.

Figures for FLFP and the incidence of part-time workers among employed women have been retrieved from OECD (2010) statistics. The labor force participation values are for women between the ages 15 and 64. Since the relationship between FLFP and elasticity size may be nonlinear, we control for FLFP using categorical variables for ranges 60–65, 65–70, and 70 or more. There are a few years missing in the data for incidence of part-time work in various countries; for these the closest possible year is used instead. Interpolating for part-time is avoided since it correlates and varies with business cycles (Buddelmeyer *et al.*, 2004). The degree of income inequality is measured by Gini coefficients. Data are retrieved from an

Table 2. Summary Statistics.

	Mean	Std. dev.	Minimum	Maximum
Elasticity	−0.259	0.252	−0.984	0
Single	0.279	0.454	0	1
Low income	0.047	0.213	0	1
Median child-age	4.419	2.101	1	9
Natural experiment	0.093	0.294	0	1
Multinomial logit	0.303	0.465	0	1
SSCI journal	0.814	0.394	0	1
FLFP 60–65	0.209	0.412	0	1
FLFP 65–70	0.558	0.502	0	1
FLFP 70+	0.140	0.351	0	1
Part-time incidence	25.808	10.205	4.7	60.2
Gini coefficient	35.528	3.64	25.8	40.8
Square root of the sample size	98.583	141.090	12.884	826.723
Sample year	1991.948	6.486	1979.392	2006

Note: The minimum value of the sample year is the weighted mean of the sample used by Cascio (2009).

updated version of the dataset compiled by Deininger and Squire (1996), along with more recent figures from the World Bank Indicators and CIA's the World Factbook (2011). Again, several years of data were missing, and data from the closest year available were used instead in these cases. In the case of studies with several years of data, we took the average of the years or used the median year. Most child care studies use one or two years of data and the variation in aggregate variables tends to be limited. One exception is Cascio (2009), who uses five separate waves of data from 1950, 1960, 1970, 1980, and 1990. In that case, we weighted year, labor force participation, and Gini coefficients variables for each year with the sample sizes of the treatment groups used for married and single women's employment estimation.[6] We could not take the weighted mean for part-time work since there are no data for the earlier years. Instead, we used the value of the year closest to the median year of 1970, which was 1979. There does not seem to be much variation in part-time work over the years, and the coefficients on the part-time variable do not change significantly when the Cascio (2009) study is excluded. A further complication arose with the studies of Michalopoulos and Robins (2000) and Michalopoulos and Robins (2002) who use data from both the United States and Canada, but we simply took the average values for macro-level variables in this case. When these two studies are excluded from the analysis, the effects of the US dummy in the meta-regressions remain similar.

Results for four separate weighted least square regressions based on 43 elasticity estimates are presented in Table 3. Since there is strong multicollinearity between some of the variables, we introduce the independent variables gradually. In Model 1, we control only for basic methodological differences, sample year, and studies from the United States. Model 2 corrects for publication bias by introducing the inverse of the square root of the sample size. It further includes a control for SSCI publications. Model 3 includes the aggregate level controls for FLFP, part-time work incidence, and Gini coefficient. Model 4 drops the controls for year and the United States to see whether the aggregate level controls' coefficients change due to multicollinearity.

Since the child care price elasticity of employment is generally negative, positive coefficients indicate correlation with smaller and negative coefficients with larger estimates. The models do not suggest a significant publication bias. The inverse of the square root of the sample size has statistically insignificant

Table 3. Determinants of Child Care Price Elasticity of Employment.

	1	2	3	4
United States	0.236**	0.242**	0.006	
	(0.115)	(0.097)	(0.117)	
Year	0.017**	0.017***	0.011	
	(0.006)	(0.005)	(0.007)	
Poverty	−0.17	−0.106	−0.004	−0.027
	(0.124)	(0.256)	(0.279)	(0.264)
Single	−0.245**	−0.274***	−0.113	−0.141**
	(0.105)	(0.096)	(0.096)	(0.065)
Age	0.049**	0.045	−0.018	−0.024
	(0.024)	(0.027)	(0.027)	(0.026)
Natural experiment	0.178*	0.113	−0.098	−0.1
	(0.092)	(0.095)	(0.081)	(0.066)
Multinomial logit	0.095	0.142	0.088	0.093
	(0.116)	(0.117)	(0.101)	(0.105)
Sample size		−3.329	−3.8	−3.588
		(6.761)	(5.195)	(4.521)
SSCI		0.259***	0.216***	0.202***
		(0.091)	(0.068)	(0.069)
FLFP 60–65			−0.311***	−0.319***
			(0.098)	(0.073)
FLFP 65–70			−0.492***	−0.412***
			(0.108)	(0.078)
FLFP 70+			−0.154	0.045
			(0.230)	(0.073)
Part-time (%)			0.004	0.009**
			(0.005)	(0.004)
Gini coefficient			−0.008	−0.006
			(0.016)	(0.016)
Observations	43	43	43	43
R-squared	0.594	0.659	0.86	0.848

Notes: ***$p < 0.01$, **$p < 0.05$, *$p < 0.1$. Robust standard errors in parenthesis. The sample size variable is the inverse of the square root of the sample size. The models are estimated using WLS with the studies' square root of the sample size as weights.

effects, while the SSCI coefficient suggests that studies published in SSCI journals report elasticities closer to 0. Given that the effect of the natural experiment variable becomes insignificant once a control for sample size is added to the model, we can conclude that the smaller estimates from natural experiments with large samples drive the positive correlation between the sample size and the elasticity size. The methodological and sample characteristics included in the regressions have lower precision in the estimates because many of them are based on a very small number of observations. None of these parameters appear to be statistically significant in all models. The low income and single parent sample controls have negative coefficients in all models, which is consistent with literature predictions.[7] The coefficient for single parent samples is particularly large. Child age has negligible effects. Multinomial logit models are correlated with smaller elasticity estimates with large coefficients of up to 0.13, but none of the estimates are statistically significant. Since there are only four natural experiments in the sample (9.3% of the total sample), the coefficient for natural experiment estimates is imprecise. The coefficient even turns negative once macro-level variables are added to the models.

The FLFP variables appear to be statistically significant and large. The difference between countries with an FLFP rate below 60% and between 65% and 70% is more than 1.5 times the mean elasticity size of −0.26. The relationship between the elasticity size and FLFP rate appears to be negative, but concave. For the estimates with FLFP rate above 70%, the negative effect becomes statistically insignificant and turns positive in Model 4. Many developed countries are now near or have surpassed the level of FLFP where price elasticity peaks. To put the results into context, according to OECD statistics, the FLFP rate among OECD countries was at 61.5% in 2009, while the corresponding value was 65.8% for EU-15 countries.

It is possible to argue for reverse causality based on the coefficients of FLFP because elasticity values may imply that governments can take advantage of high employment elasticity by increasing child care subsidies. However, the FLFP figure used here is for the entire working age population of women rather than only women in an age group with high fertility who are most likely to be affected by changes in child care prices. Furthermore, while the reverse causality argument could be plausible for the positive effect found, it is not so for the diminishing effects. If any quadratic effects were expected at all, the prediction would be to have a convex relationship, such as that of a usual cost function, between the FLFP rate and elasticity size.

Of the two other macro-level issues that were discussed in Section 2, part-time work appears to be correlated with smaller elasticity sizes. However, the coefficient is only statistically significant if sample year and the US dummy are excluded from the model. On the other hand, the Gini coefficient does not have a significant relationship with elasticity size. As a result, the difference in inequality receives little support to be counted as a major explanatory factor for the different elasticity sizes found in Europe and the United States. The very high participation rates in countries like the Netherlands, Sweden, and Norway as well as differences in the incidence of part-time employment may be more likely explanations for the differences in findings from different settings.

To test the sensitivity of the results, several alternative specifications were estimated. First, we used OLS rather than WLS to estimate the models. The coefficients remain similar, but the standard errors become larger for the FLFP variables. If we include a linear FLFP variable along with its square instead of categories, we still find concave effects and a joint significant level of 5% in OLS models. As a second test, we checked the consistency of the standard errors by clustering them at the study level. Since most studies in our sample have 1 or at most 2 estimates from them, this does not seem to affect the results as we already use robust standard errors. Finally, we used the sample size as the weighting variable rather than the square root of the sample size. While most coefficients appear to have smaller standard errors in that case, the direction of the coefficients does not change.

5. Discussion

There is a consensus that child care subsidies have positive effects on female employment, despite some recent studies showing smaller or insignificant effect sizes in various countries. This has led to a widespread view of child care subsidies as a strong policy tool for increasing female employment. However, comparison across studies is made difficult due to differences in methodological choices and macro-level factors. While the European (and Canadian) literature shows a mean employment elasticity with regard to child care prices of about −0.15, the mean elasticity of the United States only sub-sample is −0.35. The underlying reasons for these differences across countries could help to give a better understanding of what is being reported from micro-level research.

While tentative, the review of the empirical literature and the meta-regression of Section 4 show that some of the variation in elasticity sizes can be partially explained through methodological differences and macro-level factors. The elasticity estimates have become smaller over time, which is likely to be partially due to methodological improvements. However, our meta-regressions do not capture robust effects from methodological and sample characteristics. Statistically significant relationships are found between elasticity sizes and labor market characteristics. The FLFP rate has the strongest relationship with elasticity size, which is positive yet diminishing at high levels. There is also a negative, though weaker, relationship between elasticity size and part-time incidence. For countries that have reached very high participation or high part-time rates like Sweden, Norway, or the Netherlands, further policy focus on child care prices might be less effective than in the past from an employment perspective. Similarly, directly borrowing high participation countries' family policies may not dramatically increase employment in countries with very low participation rates. Considering alternatives to prices, such as quality of care offered, could help induce untapped participation effects. Already, quality of care has been examined in terms of child care demand and supply (Blau and Hagy, 1998), but its links to labor supply need further analysis.

The hypothesis that effects may be smaller in some settings does not invalidate the use of child care subsidies for other policy goals. Evidence from Germany, for which studies generally find low employment elasticities, shows that child care subsidies can improve fertility (Haan and Wrohlich, 2011). Furthermore, there is evidence that subsidized child care can have a positive effect on child development and social mobility (Havnes and Mogstad, 2011, 2015).

The review of the empirical literature on child care prices and labor supply highlights two areas for further research. Hours elasticity estimates are rare even though the intensive margin is becoming more important due to higher female participation rates in most developed countries. Second, other aspects of child care such as quality and availability are areas in which more research is needed, especially given that quality or flexibility in child care services may involve trade-offs with the price of care.

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Notes

1. The literature uses the terms participation and employment interchangeably. We use the term employment to indicate that a mother is working and female labor force participation (FLFP) to indicate aggregate participation in the labor market including unemployment.
2. While the initial database construction was done by the authors, a student assistant checked through most of the papers in our database to correct any remaining errors in elasticity, sample size, and sample year variables following the guidelines in Stanley *et al.* (2013). The authors resolved coding differences that arose during the check by referring to the primary studies and subsequent studies that cited these primary studies.
3. We estimate the WLS regressions using the *aweight* routine in Stata. Since we proxy the standard error using the square root of the sample size, the sample size is used as the proxy for the inverse variance.
4. Michalopoulos and Robins (2000) and Michalopoulos and Robins (2002) use both Canadian and American data. We categorized their estimates as outside of the United States.
5. We additionally tried to control for estimates from samples of married or single women separately but the results are contradictory and no discernible pattern emerges. Studies that include both single and married women are largely made up of married women.

6. We acquired labor force participation rate for the United States in 1950 from St. Louis FED FRED (2014) since OECD data starts in 1960. FRED (Federal Reserve Economic Data) seems to use a slightly different definition and reports lower values in later years compared to OECD statistics. The difference is most likely because FRED data report rates for all women instead of women between 15 and 64.
7. The coefficients for low income are also negative in all models and become statistically significant in the first model if Kimmel (1995), for which we do not have the exact sample size, is excluded from the model.

References

Anderson, P.M. and Levine, P.B. (1999) Child care and mothers' employment decisions. Tech. rep., National Bureau of Economic Research, No. 7058.

Averett, S.L., Peters, H.E. and Waldman, D.M. (1997) Tax credits, labor supply, and child care. *Review of Economics and Statistics* 79: 125–135.

Baker, M., Gruber, J. and Milligan, K. (2005) Universal childcare, maternal labor supply, and family well-being. Tech. rep., National Bureau of Economic Research, No. 11832.

Baker, M., Gruber, J. and Milligan, K. (2008) Universal child care, maternal labor supply, and family well-being. *Journal of Political Economy* 116: 709–745.

Baum, C.L. (2002) A dynamic analysis of the effect of child care costs on the work decisions of low-income mothers with infants. *Demography* 39: 139–164.

Bettendorf, L.J., Jongen, E.L. and Muller, P. (2015) Childcare subsidies and labour supply—Evidence from a large Dutch reform. *Labour Economics* 36: 112–123.

Blau, D. and Currie, J. (2006) Pre-school, day care, and after-school care: Who's minding the kids? *Handbook of the Economics of Education* 2: 1163–1278.

Blau, D.M. and Hagy, A.P. (1998) The demand for quality in child care. *Journal of Political Economy* 106: 104–146.

Blau, D.M. and Robins, P.K. (1988) Child-care costs and family labor supply. *The Review of Economics and Statistics* 70: 374–381.

Blau, D.M. and Robins, P.K. (1991) Child care demand and labor supply of young mothers over time. *Demography* 28: 333–351.

Borck, R. (2014) Adieu Rabenmutter—Culture, fertility, female labour supply, the gender wage gap and childcare. *Journal of Population Economics* 27: 739–765.

Buddelmeyer, H., Mourre, G. and Ward-Warmedinger, M. (2004) Part-time work in EU countries: Labour market mobility, entry and exit. Tech. rep., IZA Discussion Paper, No.1415.

Card, D., Kluve, J. and Weber, A. (2010) Active labour market policy evaluations: A meta-analysis. *The Economic Journal* 120: F452–F477.

Cascio, E.U. (2009) Maternal labor supply and the introduction of kindergartens into American public schools. *Journal of Human Resources* 44: 140–170.

CIA (2011) *CIA World Factbook*. United States: CIA, Washington DC.

Cleveland, G., Gunderson, M. and Hyatt, D. (1996) Child care costs and the employment decision of women: Canadian evidence. *Canadian Journal of Economics* 29: 132–151.

Connelly, R. (1992) The effect of child care costs on married women's labor force participation. *The Review of Economics and Statistics* 74: 83–90.

Connelly, R. and Kimmel, J. (2003a) Marital status and full–time/part–time work status in child care choices. *Applied Economics* 35: 761–777.

Connelly, R. and Kimmel, J. (2003b) The effect of child care costs on the employment and welfare recipiency of single mothers. *Southern Economic Journal* 69: 498–519.

Deininger, K. and Squire, L. (1996) A new data set measuring income inequality. *The World Bank Economic Review* 10: 565–591.

Del Boca, D., Locatelli, M. and Vuri, D. (2004) Child care choices by Italian households. Tech. rep., IZA Discussion paper series No. 983.

Doiron, D. and Kalb, G. (2005) Demands for child care and household labour supply in Australia. *Economic Record* 81: 215–236.

FRED (2014) *US. Bureau of Labor Statistics, Civilian Labor Force Participation Rate: Women*. FRED, Federal Reserve Bank of St. Louis, St. Louis.

Gong, X., Breunig, R. and King, A. (2010) How responsive is female labour supply to child care costs: New Australian estimates. Tech. rep., IZA Discussion Paper No. 5119.

Gustafsson, S. and Stafford, F. (1992) Child care subsidies and labor supply in Sweden. *Journal of Human Resources* 27: 204–230.

Haan, P. and Wrohlich, K. (2011) Can child care policy encourage employment and fertility? Evidence from a structural model. *Labour Economics* 18: 498–512.

Han, W. and Waldfogel, J. (2001) Child care costs and women's employment: A comparison of single and married mothers with pre-school-aged children. *Social Science Quarterly* 82: 552–568.

Havnes, T. and Mogstad, M. (2011) No child left behind: Subsidized child care and children's long-run outcomes. *American Economic Journal: Economic Policy* 3: 97–129.

Havnes, T. and Mogstad, M. (2015) Is universal child care leveling the playing field? *Journal of Public Economics* 127: 100–114.

Heckman, J. (1974) Shadow prices, market wages, and labor supply. *Econometrica: Journal of the Econometric Society* 42: 679–694.

Heckman, J.J. (1979) Sample selection bias as a specification error. *Econometrica: Journal of the Econometric Society* 47: 153–161.

Heckman, J. and Singer, B. (1984) A method for minimizing the impact of distributional assumptions in econometric models for duration data. *Econometrica: Journal of the Econometric Society* 52: 271–320.

Herbst, C.M. (2010) The labor supply effects of child care costs and wages in the presence of subsidies and the earned income tax credit. *Review of Economics of the Household* 8: 199–230.

Hotz, V.J. and Kilburn, M.R. (1991) *The Demand for Child Care and Child Care Costs: Should We Ignore Families with Non-Working Mothers?* Economic Research Center/NORC, Chicago.

Jenkins, S.P. and Symons, E.J. (2001) Child care costs and lone mothers' employment rates: UK evidence. *The Manchester School* 69: 121–147.

Kimmel, J. (1995) The effectiveness of child-care subsidies in encouraging the welfare-to-work transition of low-income single mothers. *American Economic Review* 85: 271–275.

Kimmel, J. (1998) Child care costs as a barrier to employment for single and married mothers. *Review of Economics and Statistics* 80: 287–299.

Kornstad, T. and Thoresen, T.O. (2007) A discrete choice model for labor supply and childcare. *Journal of Population Economics* 20: 781–803.

Lippe, T. and Siegers, J.J. (1994) Division of household and paid labour between partners: Effects of relative wage rates and social norms. *Kyklos* 47: 109–136.

Lokshin, M. (2004) Household childcare choices and women's work behavior in Russia. *Journal of Human Resources* 39: 1094–1115.

Lokshin, M. and Fong, M. (2006) Women's labour force participation and child care in Romania. *The Journal of Development Studies* 42: 90–109.

Lundin, D., Mörk, E. and Öckert, B. (2008) How far can reduced childcare prices push female labour supply? *Labour Economics* 15: 647–659.

Michalopoulos, C. and Robins, P.K. (2000) Employment and child-care choices in Canada and the United States. *Canadian Journal of Economics/Revue canadienne d'économique* 33: 435–470.

Michalopoulos, C. and Robins, P.K. (2002) Employment and child-care choices of single-parent families in Canada and the United States. *Journal of population Economics* 15: 465–493.

Nelson, J.P. and Kennedy, P.E. (2009) The use (and abuse) of meta-analysis in environmental and natural resource economics: An assessment. *Environmental and Resource Economics* 42: 345–377.

OECD (2010) *OECD Statistics Database*. Paris: Organization for Economic Cooperation and Development.

Powell, L.M. (1997) The impact of child care costs on the labour supply of married mothers: Evidence from Canada. *Canadian Journal of Economics* 70: 577–594.

Powell, L.M. (2002) Joint labor supply and childcare choice decisions of married mothers. *Journal of Human Resources* 37: 106–128.

Ribar, D.C. (1992) Child care and the labor supply of married women: Reduced form evidence. *Journal of Human Resources* 27: 134–165.

Ribar, D.C. (1995) A structural model of child care and the labor supply of married women. *Journal of Labor Economics* 13: 558–597.

Rusev, E. (2006) The relative effectiveness of welfare programs, earning subsidies and child care subsides as work incentives for single mothers. Tech. rep., University of North Caroline, Chapel Hill.

Stanley, T.D. (2005) Beyond publication bias. *Journal of Economic Surveys* 19: 309–345.

Stanley, T.D. and Doucouliagos, C. (2013) Neither fixed nor random: Weighted least squares meta-analysis. Working Paper 20131, Deakin University Faculty of Business and Law, School of Accounting Economics and Finance.

Stanley, T., Doucouliagos, H., Giles, M., Heckemeyer, J.H., Johnston, R.J., Laroche, P., Nelson, J.P., Paldam, M., Poot, J. and Pugh, G. (2013) Meta-analysis of economics research reporting guidelines. *Journal of Economic Surveys* 27: 390–394.

Tekin, E. (2007) Childcare subsidies, wages, and employment of single mothers. *Journal of Human Resources* 42: 453–487.

Van Gameren, E. and Ooms, I. (2009) Childcare and labor force participation in the Netherlands: The importance of attitudes and opinions. *Review of Economics of the Household* 7: 395–421.

Viitanen, T.K. (2005) Cost of childcare and female employment in the UK. *Labour* 19: 149–170.

Wetzels, C. (2005) Supply and price of childcare and female labour force participation in the Netherlands. *Labour* 19: 171–209.

Wrohlich, K. (2004) Child care costs and mothers? Labor supply: An empirical analysis for Germany. Tech. rep., DIW-Diskussionspapiere No. 412.

Wrohlich, K. (2006) Labor supply and child care choices in a rationed child care market. Tech. rep., IZA Discussion Paper No. 2053.