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Childcare Prices and Maternal Employment: a Meta-Analysis

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Childcare Prices and Maternal Employment:

a Meta-Analysis

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Childcare Prices and Maternal Employment: a Meta-Analysis

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Abstract

The literature estimates for labor force participation elasticity with regards to child care prices are extensive and varying. While some estimates imply substantial gains from child care subsidies, others find insignificant effects. Determining the reasons for the variance in the results and the settings in which elasticity estimates are smaller or larger is of substantial policy interest. To that end, 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 regards 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. Elasticity estimates have become smaller over time, which may partly be due to 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 significantly smaller.

Keywords: child care, female employment, meta-analysis, price elasticity

JEL classification: J13, J21 en H40

Childcare Prices and Maternal Employment: a Meta-Analysis

1. Introduction

Childcare 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 childcare subsidies, since the late 1980s, a sizable body of literature appeared investigating the impact of childcare 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 regards to childcare prices in the United States, there is a large variation in results when more recent studies from the US and also Europe are taken into account. These findings challenge the hypotheses relating high maternal employment with childcare subsidies. As an example: Bettendorf *et al.* (2015) find that an increase in the subsidy rate of childcare has a very small impact on aggregate labor supply in the Netherlands. A striking example of variation over time can be seen in Sweden. The elasticity value calculated by Gustafsson and Stafford (1992) is close to 0.9 while Lundin *et al.* (2008) find no significant effects from childcare prices on maternal employment.

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 are essentially exogenous to the analysis and 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.

The rest of the chapter is organized as follows. Section 2 provides a brief overview of the basic labor supply models with childcare. Section 3 discusses the literature findings and describes the methodolo-

gies 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

2.1. Basic Models of Childcare and Labor Supply

The importance of childcare prices for female labor supply was recognized by Heckman (1974) which is called the ‘pioneering study in the field’ by Blau and Robins (1988). The most commonly found theoretical basis for analyzing the effect of childcare prices on labor supply in the literature is grounded on the simple static labor supply model with modifications to include childcare prices and choices. A number of variants of the static model have been proposed with varying degrees of flexibility by including childcare 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 childcare is a utility function consisting of consumption C and leisure l . The maximization problem has two constraints; time $l + h = T$ and budget $C = y + (w - p)h$ where y is non-labor income, h is hours of work and p is price of formal childcare. In this basic budget constraint, the price of childcare is treated like a tax on wages with each hour of work requiring the purchase of an hour of childcare. The implications of changes in the price of childcare 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 childcare 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 childcare 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 childcare. 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 childcare quality is equal across care types, the price of childcare 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 childcare 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:

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

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

$$w_r \geq p + \frac{U_l + U_q Q_l}{U_c + U_q Q_c} \quad (2)$$

The equation (2) is similar to the result that Ribar (1995) derives with slight differences in the way unpaid care and childcare prices are defined. Evaluated at the corner of the budget constraint where $h = 0$, the comparative statistics are self-evident. Childcare 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. First, the Google Scholar and EconLIT search engines were used for starting in February of 2011, 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 "childcare". The second source of literature were 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 USA and Canada, and the latter for studies from Europe. 44 estimates are included in the table 1 for analysis from around 37 English language articles published between 1988 and 2010. The coding in the

majority of studies was also done by a research assistant in December 2014.² 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 childcare expenditures and the samples sizes are usually below 10,000. Natural experiments are an exception because administrative and census data without micro-level information on childcare 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 correct for publication bias using a precision factor, usually the standard error. The basic premise of the correction is that precise estimates will be closer to the true effect. Therefore, if estimates' precision is significantly related to the effect size in one direction or the other, there is evidence for 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 report a standard error 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 childcare 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 square root of the sample size as our precision variable and include the same variable in the meta-regressions as a control variable following the example of Card *et al.*

(2010).

Table 1 shows the elasticity estimates across the literature. The employment elasticity 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, which might indicate publication bias towards smaller estimates. However, there are also significant differences in the sample make-ups and methodologies of studies, which might explain the differences. 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 over two articles 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 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 childcare 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 non-working mothers. Heckman (1979) 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) raises 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 childcare. The standard approach is to divide reported childcare costs with working or childcare hours to arrive at a price of childcare. The definition

Table 1: Literature summary

<i>Authors / Publishing Year</i>	<i>Elasticity</i>	<i>Sample Size</i>	<i>Year</i>	<i>Country</i>	<i>Data Source</i>	<i>Sample</i>
Doiron and Kalb (2005)	-0.02	5305	1997	Australia	CCS SIHC	Single
Doiron and Kalb (2005)	-0.075	1116	1997	Australia	CCS SIHC	Married
Gong <i>et al.</i> (2010)	-0.287	4184	2006	Australia	HILDA	All
Powell (1997) ³	-0.38	9201	1988	Canada	CNCCS	Married
Baker <i>et al.</i> (2005) ⁴	-0.236	33708	1998	Canada	NLSCY	Married
Cleveland <i>et al.</i> (1996)	-0.388	3976	1988	Canada	CNCCS	Married
Wrohlich (2004) ⁵	-0.025	1345	2002	Germany	SOEP	All
Wrohlich (2006)	-0.02	3213	2002	Germany	SOEP	Married
Del Boca <i>et al.</i> (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)	-0.872	166	1984	Sweden	HUS data	All
Lundin <i>et al.</i> (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

Table 1: Literature summary cont.

<i>Authors / Publishing Year</i>	<i>Elasticity</i>	<i>Sample Size</i>	<i>Year</i>	<i>Country</i>	<i>Data Source</i>	<i>Sample</i>
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)						
Kimmel (1995)	-0.01	2032	1986	USA	NLSY	All
Kimmel (1998)	-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
Michalopoulos and Robins (2000)	-0.156	13026	1989	USA/Canada	CNCCS	Married
Michalopoulos and Robins (2002)	-0.259	1762	1989	USA/Canada	CNCCS	Single

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 childcare utilized. Some studies such as Blau and Hagy (1998) avoid having to use user provided data through additional data from childcare providers.

Besides the structural models with two outcomes, employment or non-employment, 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 twelve 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 to hold for female labor supply and childcare. 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 childcare 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 childcare 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 do not differ from other structural studies based on childcare 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 US is quite large.

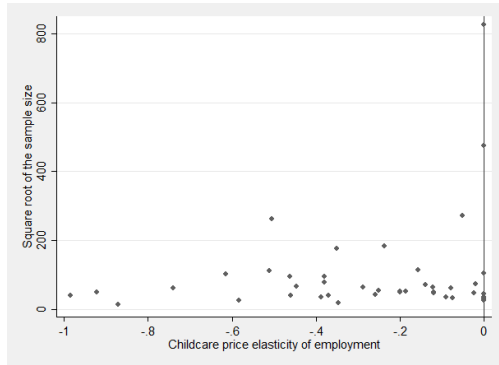


Figure 1: Sample size and elasticity estimates

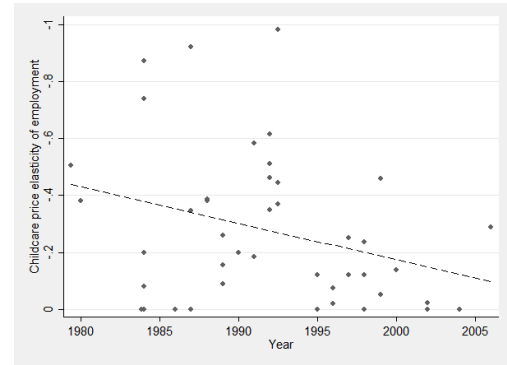


Figure 2: Elasticity estimates over time

Taking the mean of the two subsamples shows a mean of -0.19 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 childcare prices. Limiting the comparison to the United States, Tekin (2007) and Herbst (2010), who use relatively later samples from 1997 and 1999 respectively, find much smaller elasticities than estimates based on samples from the 80's or early 90's. However, it is difficult to conclude directly that the smaller elasticity findings are merely due to changes in population's responsiveness to childcare 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 childcare price elasticity and three aggregate indicators: female labor force participation, 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 female labor force participation and childcare. In one equilibrium there is little demand for childcare provision and low participation 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 childcare 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 non-maternal childcare. Countries with low female labor force participation may then also have low childcare price elasticity. However, there can also be a limit to how far childcare prices can affect participation. Lundin *et al.* (2008) find insignificant effects from childcare prices on maternal employment in Sweden and argue that further reductions in childcare 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 elas-

ticity size simply because the denominator in the elasticity calculation becomes larger. High participation figures can also signal factors influencing the individual employment decision. Countries with high female labor force participation may already have cheap and readily available formal childcare or can make use of alternative informal care arrangements, both of which will diminish the effects of further reductions in childcare prices. Furthermore, if other structural factors, such as the wage structure or working hour flexibility, contribute to high female employment, the effect of the childcare 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 childcare, while at the same time allowing women to provide informal childcare. These two effects of part-time work are expected to increase both the demand for and the supply of informal childcare which can be substituted for formal childcare. The substitution of informal care for formal care has been offered up as an explanation as to why previous literature reviews find that the price elasticity of demand for formal childcare is much larger than the labor supply elasticity with regard to the price of childcare (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 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 childcare 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 investigate the impact 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. Simulations by Stanley *et al.* (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 use the square root of the sample size as a precision factor. As previously noted, we test for publication bias through the square root of the sample size and further also add a variable indicating whether the article was published in an SSCI journal, which might be correlated with the quality of analysis. The estimated specification is given as:

$$\beta_i = X'\gamma + Z'\phi + \varepsilon_i \quad (3)$$

In equation (3), β_i is the elasticity estimate of study i , X are controls for sample and methodological characteristics including sample size, and Z represents the macro-level factors which includes a dummy for studies from the USA and the year of the sample the study uses.⁶ ε 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.⁷ 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 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, 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 female labor force participation (FLFP) and the incidence of part-time workers among em-

Table 2: Summary statistics

	Mean	Std. Dev.	Minimum	Maximum
Elasticity	-0.277	0.267	-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	0.279	0.454	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

ployed 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 female labor force participation (FLFP) and elasticity size may be non-linear, we control for FLFP using categorical variables for ranges 60 to 65, 65 to 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 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 was used instead in these cases. In case of studies with several years of data, we took the average of the years or used the median year. Most childcare 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 5 separate waves of data from 1950, 1960, 1970, 1980, 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.⁸ 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 models that include the part-time variable give similar results 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 US and Canada, but we simply took the average values for macro

Table 3: Determinants of childcare price elasticity of employment

	(1)	(2)	(3)	(4)
US	0.056 (0.092)	0.039 (0.097)	0.043 (0.093)	
Sample year	0.013* (0.006)	0.007 (0.006)	0.011 (0.008)	
Low income	-0.131 (0.148)	-0.145 (0.161)	-0.087 (0.184)	-0.098 (0.176)
Single	-0.103 (0.113)	-0.068 (0.120)	-0.024 (0.121)	-0.034 (0.117)
Child age	0.010 (0.024)	-0.006 (0.028)	-0.020 (0.027)	-0.026 (0.025)
Natural experiment	0.194* (0.098)	0 (0.128)	-0.121 (0.123)	-0.133 (0.142)
Multinomial logit	0.146* (0.086)	0.178* (0.092)	0.128 (0.107)	0.148 (0.105)
Sample size		0.041* (0.020)	0.040 (0.035)	0.043 (0.034)
SSCI		0.073 (0.097)	0.168** (0.071)	0.131* (0.077)
FLFP 60-65			-0.258** (0.094)	-0.245** (0.104)
FLFP 65-70			-0.380*** (0.095)	-0.309*** (0.092)
FLFP 70+			-0.222 (0.192)	-0.075 (0.138)
Part-time (%)			0.005 (0.004)	0.008** (0.003)
Gini coefficient			-0.007 (0.014)	-0.002 (0.013)
Observations	43	43	43	43
R-squared	0.332	0.393	0.594	0.568

*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$. Robust standard errors in parenthesis. The sample size variable is the square root of the sample size.

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 independent variables gradually. In model 1, we control only for basic methodological differences, sample year and studies from US. Model 2 introduces methodological variables. Model 3 corrects for publication bias by introducing the square root of sample size and a control for SSCI publications and adds the aggregate level controls for female labor force participation, part-time work incidence and gini coefficient. Model 4 drops the controls for year and US to see whether the aggregate level controls' coefficients change due to multicollinearity.

Since the childcare price elasticity of employment is generally negative, positive coefficients indicate correlation with smaller and negative coefficients with larger estimates. The models suggest that there

is significant publication bias. The sample size has a statistically significant effect in model 3 and has a consistently positive coefficient and a similar positive and statistically significant coefficient is found for the SSCI variable. The positive coefficients suggest that studies with larger samples sizes or those that are published in SSCI journals tend to have elasticity estimates closer to 0. Combined with the result that the impact of sample year becomes statistically insignificant once publication bias controls are included, the results suggest that the earlier studies with smaller samples may have had a bias towards reporting larger elasticity values. 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. The low income sample control has a negative coefficient, which is consistent with literature predictions⁹. Child age has negligible effects. In models 2 and 3, multinomial logit models are shown to have significantly smaller elasticity estimates with large coefficients of up to 0.2. 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 the sample size is controlled for since the sample size tends to be very large in natural experiments and has a positive effect itself.

The labor force participation variables appear to be statistically significant and large. The difference between countries with a female labor force participation rate below 60% and between 60% and 70% is nearly the mean elasticity size of 0.29. The relationship between the elasticity size and female participation rate appears to be positive, but concave. For the estimates with female labor force participation above 70%, the positive effect becomes statistically insignificant and smaller once sample year is excluded from the model. Many developed countries are now near or have surpassed the level of female labor force participation where price elasticity peaks. To put the results into context, according to OECD statistics, the female labor force participation 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 female labor force participation because elasticity values may imply that governments can take advantage of high employment elasticity by increasing childcare subsidies. However, the female participation 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 childcare prices. Furthermore, while reverse causality argument could be plausible for the positive effect found, it is not 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 participation rates and elasticity sizes.

Of the two other macro-level issues that were discussed in section 2, part time work appears to be correlated with smaller elasticity estimates. However, the coefficient is only statistically significant if sample year is excluded from the model. On the other hand, the Gini coefficient does not have 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 USA. 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 unreported robustness checks were performed. First, we used OLS rather than WLS to estimate the models. The coefficients remain similar, but the standard errors become larger for female labor force participation variables and smaller for the sample size variable. The concavity of the effects seems to disappear, as a simple linear measure of female labor force participation had significant positive effects at the 1% level 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. We also tested an alternative clustering strategy where estimates from the data source and sample year were clustered. Clustering based on the data source does not appear to change the significance levels. Finally, we used the sample size as the precision factor rather than the square root of the sample size. While some coefficients appeared to have smaller standard errors, the overall results did not change.

5. Discussion

There is a consensus that childcare 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 childcare subsidies as a rather 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 show a mean employment elasticity with regards to childcare prices of about -0.19, the mean elasticity of the US only sub-sample is -0.35. The underlying reasons for these differences 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 explained through methodological differences and macro-

level factors. Analyses based on subsamples of low income mothers are associated with larger elasticity sizes, which lends support to subsidies based on means-testing. Model choice, especially discrete choice models of varying sophistication, can similarly alter the size of the employment elasticity. More significant relationships are found between elasticity sizes and labor market contexts. The female labor force participation rate has a positive yet diminishing relationship with the elasticity size, while proportion of women in part-time work has a negative relationship with elasticity size. For countries that have reached very high participation or high part-time rates like Sweden, Norway or the Netherlands, further policy focus on childcare 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 costs, such as quality of care offered, could help induce untapped participation effects. Already, quality of care has been examined in terms of childcare 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 childcare subsidies for other policy goals. Evidence from Germany, for which studies generally find low employment elasticities, shows that childcare subsidies can improve fertility (Haan and Wrohlich, 2011). Furthermore, there is evidence that subsidized childcare can have a positive effect on child development and social mobility (Havnes and Mogstad, 2011, 2014).

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

Notes

¹The 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 to indicate aggregate participation in the labor market including unemployment.

²While the initial database construction was done by the authors, a student assistant checked through most of the papers and our database to correct any remaining errors following the guidelines in Stanley *et al.* (2013). Several mistakes in sample sizes were corrected as a result of this check. One of the authors had had previous coding experience for an unpublished meta-regression analysis.

³The 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 childcare 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.

⁴This paper was 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.

⁵The 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 6.

⁶Michalopoulos and Robins (2000) and Michalopoulos and Robins (2002) use both Canadian and American data. We categorized their estimates as outside of the US

⁷We additionally tried to control for estimates from samples of married or single women separately but the results are contradictory and no discernable pattern emerges. Studies that include both single and married women are largely made up of married women.

⁸We 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 reports rates for all women instead of women between 15 and 64.

⁹The coefficients become larger and statistically significant at the 5% level if Kimmel (1995), for which we do not have the exact sample size, is excluded from the model.

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