Institutional Setting Matters for the Effect of Childcare on Maternal Labor Supply. A Cutoff-based Cross-country Analysis

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Preliminary draft

January 3, 2017.

Abstract

Previous evidence from single country analyses suggest that estimates of the effect of (subsidized) childcare availability on maternal labor supply varies greatly by institutional context. We provide conditional estimates of the childcare effect that directly capture its interaction with other institutional factors, based on harmonized data from 6 EU countries. The identification of the effect of childcare is based on variation due to birthdate-based eligibility cutoffs specific to each country. We estimate a comprehensive model on pooled data, which includes country-level measures of key institutional factors and allows for their interactions with childcare availability measures. This method utilizes the cross-country variation in institutional characteristics to provide conditional childcare effect estimates, which sheds light on the causes of the variability in estimated childcare effects across contexts. The estimates suggest that higher benefit replacement rates, availability of part-time job opportunities and short leave length increase the childcare effect on maternal LFP. The results are also useful to policymakers in the evaluation of plans for further childcare expansion.

Keywords: subsidized childcare, maternal labor supply, institutional context

® Certified random order of the authors. (Ray and Robson, 2016)

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This research was supported by a grant from the CERGE-EI Foundation under a program of the Global Development Network. All opinions expressed are those of the authors and have not been endorsed by CERGE-EI or GDN. We are grateful to the participants of the GDN Summer Workshop for their valuable comments.

1. Introduction

Previous evidence on the effect of subsidized childcare availability on maternal labor supply suggests that the effect varies greatly among countries due to differences in their institutional and cultural contexts (Cascio et al., 2015). Yet there is little direct evidence on the interdependencies of childcare availability and other factors, and policymaking is mostly limited to general targets for childcare coverage – for example, the EU's Barcelona Targets[1] – that are not linked to reforms of other potential limiting factors of maternal labor force participation. This paper provides quasi-experimental estimates of the childcare effect based on harmonized data from 6 countries[2] (Austria, France, Hungary, Czech Republic, Slovakia and Greece), and a comprehensive model that allows us to quantify its interactions with other factors in the form of conditional childcare effect estimates. The analysis follows in the footsteps of two main strands of literature: quasi-experimental single country estimates of the childcare effect, and policy analyses based on cross-country comparisons of country-level institutional indicators and maternal outcomes. The comprehensive model is unable to provide unbiased estimates of the cross-effects of childcare and other institutions as a result of the endogeneity of the institutional variables, nevertheless, it provides new insights into how childcare effect varies with institutional context. The results provide policymakers with more accurate information regarding the potential benefits of childcare expansion, which directly takes the countries' contexts into account.

Methodologically, recent research on the childcare effect has increasingly turned towards quasi-experimental methods based on policy changes or birthdate-based eligibility cutoffs. While these allow for better identification of the childcare effect due to the exogenous source of variation, they are local in nature, and therefore highly dependent on the estimation context. The results of single country estimates vary accordingly, as they come from very different institutional settings and measure the impact at various child ages. A recent study highlights the relevant factors that most probably drive the differences (Cascio et al., 2015). First, the actual labor supply rate of mothers affected by the treatment - i.e. its level prior to treatment – is key, since the scope for policy to increase labor supply would be limited by an already high rate. Second, interdependencies with other institutional elements - such as child-related leaves, labor market flexibility, and cultural norms - are also important. Quasi-experimental evidence from various countries is in line with these points. No effect or a very small effect was found in the US (Cascio, 2009; Fitzpatrick, 2010), and France (Givord and Marbot, 2015), where maternal employment rates of the treated were already high. A more significant impact was found in Spain (Nollenberger and Rodríguez-Planas, 2015), in 1996 Germany (Bauernschuster and Schlotter, 2015), and in Hungary (Lovasz and Szabo-Morvai, 2013) in settings where pre-treatment maternal employment rates were significantly lower. Some studies qualitatively discuss the role of the

[1] The European Union set specific targets for its countries in 2002 and renewed them in the Europe 2020 Strategy, prescribing a 33% coverage rate for children under 3, and a 90% coverage rate for those between 3 and the mandatory school age by 2010 (EC, 2013, 2008). While most previous estimates pertain to western countries with relatively supportive environments and already high maternal labor supply rates, little evidence is available from settings with very different institutional contexts, such as the Southern and Central-Eastern European (CEE) countries. Since most of these countries are significantly behind in fulfilling their obligations and expansion places a high financial burden on them, it is important to assess the expected labor market impact accurately given their particular context.

[2] The current version of the paper is based on 6 countries, but we are hoping to add further countries as the data becomes available. Estonia, Italy, Latvia, Lithuania and Portugal are potential candidates.

leave system and cultural views in constraining the childcare effect (Givord and Marbot, 2015; Nollenberger and Rodríguez-Planas, 2015), and that of highly qualified mothers and the lack of childcare alternatives in magnifying it (Bauernschuster and Schlotter, 2015). The methodology used in this paper allows us to directly quantify these relationships.

While quasi-experimental studies generally focus on the effect of a single policy, another strand of policy literature analyzes the roles of a set of family policies and cultural norms on maternal labor supply based on cross-country comparisons (Boca et al., 2009; Cipollone et al., 2014). These studies generally find that the availability of childcare - especially under age 3 -, the existence of job-protection and well-paid leave that are neither too short nor too long, flexible job opportunities, and cultural support for maternal employment lead to differences in the employment participation and working hours of mothers compared to childless women (Boeckmann et al., 2014). These articles provide a solid ground for considering which institutional elements to include in the cross-country analysis of the conditional childcare effect. Direct evidence on the interactions of childcare and other factors is scarce. (Budig et al., 2012) show that cultural attitudes moderate the impact of policies on women's earnings across countries. (Cukrowska-Torzewska, 2015) studies the effect of various policy measures on maternal employment and wages, based on data from 28 European countries, allowing for the interaction of childcare availability and leave policies. The findings indicate that the impact of leaves is dependent on childcare availability: long maternity leaves combined with high childcare coverage lead to a higher gap in the employment of mothers and non-mothers compared to settings where the coverage is low. Our study focuses on the estimation of the causal effect of childcare availability, rather than the full set of policy measures, and utilizes the exogenous variation due to eligibility cutoffs for more precise identification.

This analysis combines representative harmonized European Labour Force Survey (EU LFS) data from 6 countries (covering 2005-2012) with country-level information about kindergarten enrollment cutoffs (provided by country experts and conformed by data) and country-specific institutional characteristics based on various sources of information like Multilinks database, OECD Family Database, European Social Survey, etc. As a first step, we document country-level differences in the timing of maternal labor market return process after having a child, relative to major changes in family policies and other institutional factors. We then estimate the country-specific effects of childcare on maternal labor supply using an instrumental variables (IV) method based on childcare eligibility cutoffs around age 3 of children, and compare them in light of further characteristics of the institutional environments. Finally, we propose a comprehensive framework that recognizes the importance of examining the policy mix as a whole, but retains the cutoff-based strategy for the identification of the childcare effect. We use the pooled country dataset to estimate the conditional effects of childcare availability and certain elements of the institutional framework (leaves, flexible work, societal attitudes) on maternal labor supply.

There are two important novelties in this paper. The first is using cross-country cutoff based estimates in order to uncover the interaction between childcare and other relevant institutions. This task is rather difficult, as the institutions (rules, policies, norms etc.) historically evolve together with labor market outcomes and thus are endogenous. As a result the estimated coefficients of their interactions are likely biased. Still, the results highlight how childcare effect varies with institutional context. The results indicate that higher benefit replacement rates, availability of part-time job opportunities and short leave length increase the childcare effect on maternal LFP. The second is that the paper provides

cutoff-based childcare effect estimates for some countries from which such measures were not available previously (like Greece and Slovakia). The single-country estimates are comparable due to the harmonized data and the unified estimation method. The results from the pooled regressions suggest that child benefits weaken the childcare effect, whereas norms amplify that.

2. Institutional context

Our sample of countries is determined by data availability and the existence of kindergarten eligibility cutoffs that are necessary for the IV-based analysis. The countries included differ significantly in terms of their institutional environments, family policies, and cultural norms. Table 1 summarizes characteristics that most probably play an important role in shaping maternal labor supply for the countries in our analysis. The choice of institutional measures builds on previous empirical findings of the policy literature. Previous studies suggest that too short leaves (or non-existent) may constrain the opportunities of women to reenter their jobs, and discourage women from higher income households to return to work. On the other hand, long, low-paying leaves may lead to (especially low-skilled) mothers becoming detached from the labor market and the depreciation of their skills, as well as increased statistical discrimination against mothers and women (Boeckmann et al., 2014). We therefore focus on two measures of leave policies: the length of paid leave (job protection) available to mothers, and the amount of the benefit that is available to mothers at the child age analyzed in our study.[3] The flexibility of labor markets is also an important factor, though it lies outside the direct realm of family policies. The employment rate of mothers with young children is strongly correlated with the availability of part-time work opportunities: part-time work may provide mothers with a means to strengthen their attachment to the labor market and keep their skills up to date, while allowing for a more gradual separation from their child. The quality (related job protection, social benefits and earnings) of the available part-time jobs may also matter (Del Boca, 2002). In the analysis we include the ratio of part-time work within all jobs as a proxy for flexibility. Cultural norms are also strongly correlated with maternal outcomes, and unfavorable attitudes may limit the effectiveness of family policies (Budig et al., 2012).[4] Based on European Values Survey data, we use country-specific measures of norms and attitudes.

Table 1: Institutional characteristics of the countries

Country	Childcare enrollment at age 2	Childcare enrollment at age 3	Total paid leave length (weeks)	Total leave - average replacement rate	Cash benefits at age 3	Share of female part-time in employment	Norms (preschool child suffers if mother works)
Austria	17%	60%	60	85	12%	73%	56%
France	51%	98%	42	45	11%	36%	34%
Hungary	9%	74%	160	44	59%	11%	52%
Czech Republic	6%	60%	110	51	46%	11%	46%
Slovakia	5%	60%	164	32	39%	6%	41%
Greece	21%	67%	43	54	5%	9%	62%

[3] Further potential leave measures: job protection, overall benefit amount/replacement rate, length of leave for fathers. We plan to include these alternative measures later on. Another aspect of the design of leaves is how much they enable and encourage fathers to take part in childcare duties. Among western countries, those that explicitly encourage the more even distribution of duties among mothers and fathers have the highest levels of maternal and female employment, though the gender gap in leave usage remains significant even in the most degenderized countries (Thévenon and Solaz, 2013).

[4] The 2010 EC report on the evaluation of the fulfillment and effectiveness of the Barcelona childcare targets also notes the importance of norms related to parenthood, institutionalized childcare, and parental preferences at the country level, and the need for these norms to be shaped through raising public awareness (Mills et al., 2014).

Table 1 reveals significant differences in maternal employment-related characteristics among the countries included in our sample. Childcare enrollment rates reveal significant variation, especially for children at age 2. France represents the western countries with rates well above the other countries at both age 2 and 3. The former socialist CEE have very low rates at age 2 of under 10 percent, and Austria and Greece lie in the middle. In terms of leaves, the CEE countries have very long, low paying leaves, while the other three countries have shorter, mostly better paid leaves. Part-time job opportunities are scarce in the CEE countries and Greece. In terms of norms regarding maternal employment with a small child, France shows the most favorable views, while Austria and Greece are the most traditional. Overall, we can say that the three CEE countries are similar in these characteristics likely due to their shared socialist heritage, with its legacy of institutions that do not support maternal employment, and views that are less supportive than the most liberal western countries, but more supportive than the traditional European countries. France represents the more liberalized western countries, with institutions and norms that favor maternal employment. Austria and Greece are more difficult to categorize, being similar to western countries in some characteristics, yet less supportive than CEE countries in others (norms). Appendix Figure A1 depicts the rate of maternal employment by child age for each country. It shows that the dynamics of mothers' return to the labor market as a function of the age of the youngest child is rather dispersed between the countries. The likely reason for this is the fact that the countries differ in many relevant aspects, as seen in Table 1. The three CEE countries show the lowest rates under age 3 of children – in line with institutions that do not support employment under age 3 – but higher rates than the traditional countries (Austria, Greece) at older child ages. Maternal employment is already significantly higher than the other countries under age 3, in line with its supportive institutions. We utilize these differences between the countries to find out how the context determines the effect of childcare availability on maternal labor supply.

3. Data and Methodology

3.1. Data

The analysis is based on individual-level EU LFS data from 6 countries. The sample of countries is determined by (a) the existence of a kindergarten eligibility cutoff, (b) the availability of birth date and age information of the youngest child of each mother in the EU LFS dataset, and (c) the availability of the institutional measures used to describe the context. To determine which countries have a birthdate-based eligibility cutoff in effect, we surveyed experts from each potential country, asking for detailed information regarding kindergarten enrollment rules and their changes over time. We then verified the cutoffs using EU-SILC childcare enrollment and birth quarter data, by comparing enrollment rate means by birthdate groups over various child ages. The cutoff of the birthdate varies by country, while the enrollment date is generally September 1st, so this is when treatment occurs. This means that in different countries, the age when treated children actually enroll varies. We summarize the details of each cutoff by country in Table 4.

The EU LFS dataset contains the exact age of the youngest child in each household up to age 14. For the countries included in the analysis we constructed a stochastic panel based on the EU LFS dataset.[5]

[5] We utilize a linking procedure to link household observations over time, for each country where the data was originally collected as a panel dataset. Linking is based on exact matches (or logical increases/decreases) of 56 variables describing the household level characteristics, household composition, and individual characteristics of certain members of the household.

We derive the birth month of the youngest child by observing in which wave their age changes.[6] We then include the mothers of the youngest children in each household in our dataset, and observe their labor force participation, employment, as well as individual and household level characteristics. Table 2 depicts some descriptive statistics of the resulting dataset.

Table 2: Descriptive statistics of the sample by country (2005-2012)

Country	# of observations	LFP (Control)	LFP (Treatment)
Austria	1,568	75%	80%
Czech Republic	1,814	23%	37%
France	4,805	77%	85%
Greece	2,481	67%	68%
Hungary	3,066	34%	53%
Slovakia	1,267	37%	59%

Finally, we link country-level institutional variables to the mother-level data gathered from various sources. In our baseline specification, we include measures of the total length of paid leave available to mothers and the level of cash benefits accessible at the child age studied (OECD Family Database), the ratio of part-time jobs to all jobs of females, and a rudimentary measure of the public attitude towards gender roles and maternal employment, specifically, the importance of family relative to work, as detailed in Table 3.

Table 3: Description of the institutional variables

Variable	Description	Data source	Data structure		
			By country	By year	By age of youngest child
Child benefit (%)	Family cash benefit as a proportion of median working-age household income	OECD Family Database	x	x	x
Part-time jobs (%)	Part-time employment of females (20-50)	EU-LFS	x	x	
Leave length (yr)	Length of paid leave (years after birth)	OECD Family Database	x		
Family > Work	Importance of family relative to work (20-50 year old females) *1 = Family is very important, work is not at all important* *.5 = Both are equally important*	European Values Study	x	x	

[6] A given household is observed every 3 months for 6 times in a row (or less in case of attrition). In each month there are certain number of households observed. Thus households observed in every first month of the quarters are different from those observed in the second and the third quarters. The exact day and month of birth of the youngest child is eliminated from the database for data security reasons. We thus only have the age of the child. However, if we have at least 4 quarters of observations in a row, we can infer to the quarter of birth from observing the age jumping up by one from one wave to the other. From that we know that the birthdate of the child should be between the two interview months. We assign the month of the latter interview to the child as being the quasi month of birth, but we only know that the month of birth was that particular month, or one of the two previous months. We have such quasi birth months in each month of the year exactly because there are interviews happening in each month. As a result, the birth data are known to a quarterly precision, and we have birth data with a monthly frequency. We take this into consideration when determining our treatment and control groups around the eligibility cutoff by excluding the 3-month birth date groups overlapping treatment and control birth periods.

	0 = Family is not at all important, work is very important				

3.2. Empirical specification

We first estimate the childcare effect for each country, based on a cutoff-based IV methodology similar to what was used in the case of Hungary (Lovasz and Szabo-Morvai, 2013). We then develop a comprehensive pooled country model to quantify the interactions of childcare and other factors.

3.2.1. Cutoff-based estimates of the childcare effect by country

The basic idea of the cutoff-based methodology, inspired by Angrist and Krueger, is to use the birthdate of the child to sort mothers into treatment and control groups (Angrist and Krueger, 1991). We compare the labor supply of mothers on the two sides of the cutoff when their children are of the age at which kindergarten generally begins for the treatment group. Since kindergarten availability is significantly higher than nursery availability (see Table 4), this means that treatment mothers have a significantly higher probability of being able to enroll their children in childcare compared to control mothers. The treatment variable is defined as follows:

$$T_i = \begin{cases} 1 & \text{if} \quad cutoff\ date - 5\ months \leq b_i \leq cutoff\ date \\ 0 & \text{if} \quad cutoff\ date \leq b_i \leq cutoff\ date + 5\ months \end{cases} \qquad (1)$$

where b_i is the date of the birthday of the youngest child, and the cutoff date varies by country (see Table 4).

In order for the estimated treatment effect to be unbiased, we need sorting into treatment to be random so that the groups differ only in terms of treatment status. By the standard argument of the regression-discontinuity design, the selection of mothers into the groups can be regarded random if the window around the cutoff is narrow enough: mothers of children born on December 31 can be assumed to be very similar to mothers of children born on January 1. The wider windows of 5 months around the cutoff, which are needed to ensure a large enough number of observations, mean that we need to consider certain possible sources of bias more carefully. Other age-related changes can lead to significant differences between the groups, because the average age of children in the two groups differs significantly.[7] In order to separate these other effects from the childcare effect, we define the estimation sample so that we include mothers in the treatment and control groups when their children are the same age. This sampling design ensures that child age, and therefore any further age-related characteristics - for example, child development or preferences regarding separation from the child - will be the same on average in the two groups.

To estimate the causal effect, we turn to IV estimation, where treatment (T) is an instrument for childcare availability. We estimate reduced form regressions separately for each country of the following form:

$$LFP_{yi} = \beta T_{yi} + \alpha_y + X'_{yi}\pi + \xi_{yi} \qquad (2)$$

where subscripts indicate yearly (y), and individual (i) variation, and LFP_{yi} is the labor force participation (LFP) dummy for individual i. The equation adjusts for a set of individual (X_{yi}), α_y represents year fixed effects. The parameter β captures the effect of belonging to the treatment group on the LFP probability. It can be interpreted as representing how much more active mothers are if they are eligible for kindergarten rather than nursery school, which has significantly lower coverage. Since

[7] With 5 month windows, child age differs by an average of 5 months between the two groups at any point in time, so the effects of these differences may be significant. Preferences regarding separation from the child are also likely to change significantly during these 5 months around age 3 (Blaskó, 2011).

these rates differ by country, the magnitude of the childcare effect estimates need to be interpreted based on their mean differences.

3.2.2. Pooled estimates of the childcare effect

We next turn to a more comprehensive analysis based on a pooled dataset from all countries in the sample. In the pooled database the cutoff date and the enrollment age of the treated varies by country. The treatment and the control groups are always defined by the birth dates of the children relative to the cutoff date of the specific country. The sample consists of children of the same age in relative terms, compared to the enrollment age of the treated. For instance, in Slovakia the enrollment age of the treated is 3 years, whereas the same in Austria is 2.5. (see Table A1 in the Appendix for details.) If we select the sample so that children of relative age of .75 years are included, this results that 3.75-year-olds are included in the Slovakian treatment and control group, and 3.25-year-olds in the Austrian. The reason is that we want to observe mothers across countries with the same length of time elapsed since enrollment, which is the beginning of the treatment. The dependent variables are LFP rates of mothers in the sample. The following regression is estimated on the individual data:

$$LFP_{yci} = \beta_1 T_{yci} + \alpha_y + \beta_2 I_{yc} + \beta_3{}' T_{yci} * I_{yc} + \pi' X_{yci} + \xi_{ycgi} \qquad (3)$$

where the estimated childcare effect is given by β_1 and the elements of the vector β_3, the coefficients of the institutional interaction terms. The subscripts indicate variation by year (y), country (c), group (g) and individual level variation. The equation adjusts for a set of demographic covariates (X_{yci}), such as the highest level of education, marital status, number of children in the household and the number of inactive persons in the household to capture the possibilities for private childcare arrangements. α_y represents year fixed effects and I_{yc} is a vector of institutional factors in the given year and country, relevant to individual i.

4. Results

First, a set of graphs are presented (Figs. 1, 2 and 3) showing different outcomes around the cutoff point. These graphs are produced using the pooled database, where the cutoff date and the enrollment age of the treated are specific to each country. The ages referred to by the graphs are relative ages compared to the enrollment age of the treated. To the left of the cutoff point are the treated mothers and to the right are the controls. As can be seen, there are rather large differences around the cutoff dates at relative age .75. At this age the children in the treatment group have already been enrolled for .75 years and their mothers had enough time to adjust their labor market status to the new situation of having access to adequate childcare; whereas children in the control group stand right before enrollment.

Figure 1: LFP rates of mothers with children born before and after the cutoff date

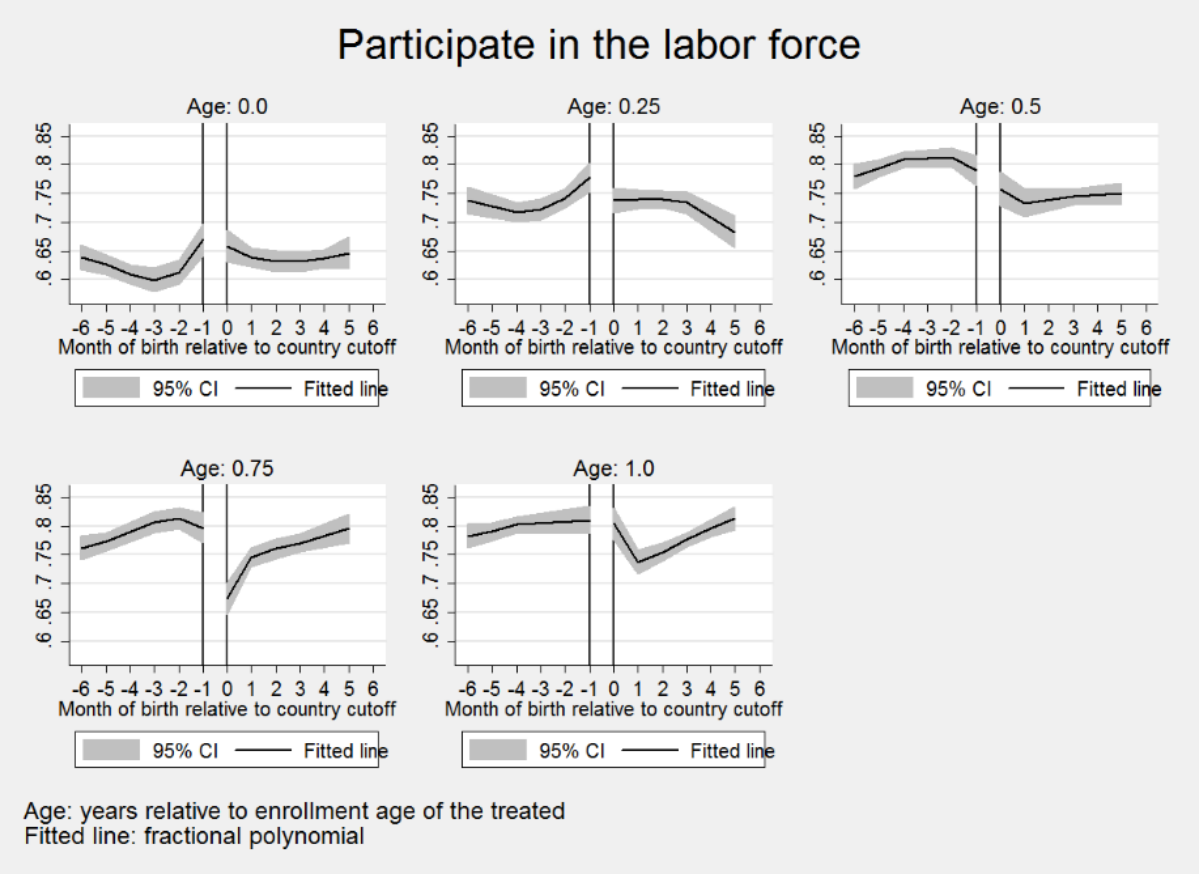


Figure 2: Reason for not looking for a job (mothers with children born before and after the cutoff date)

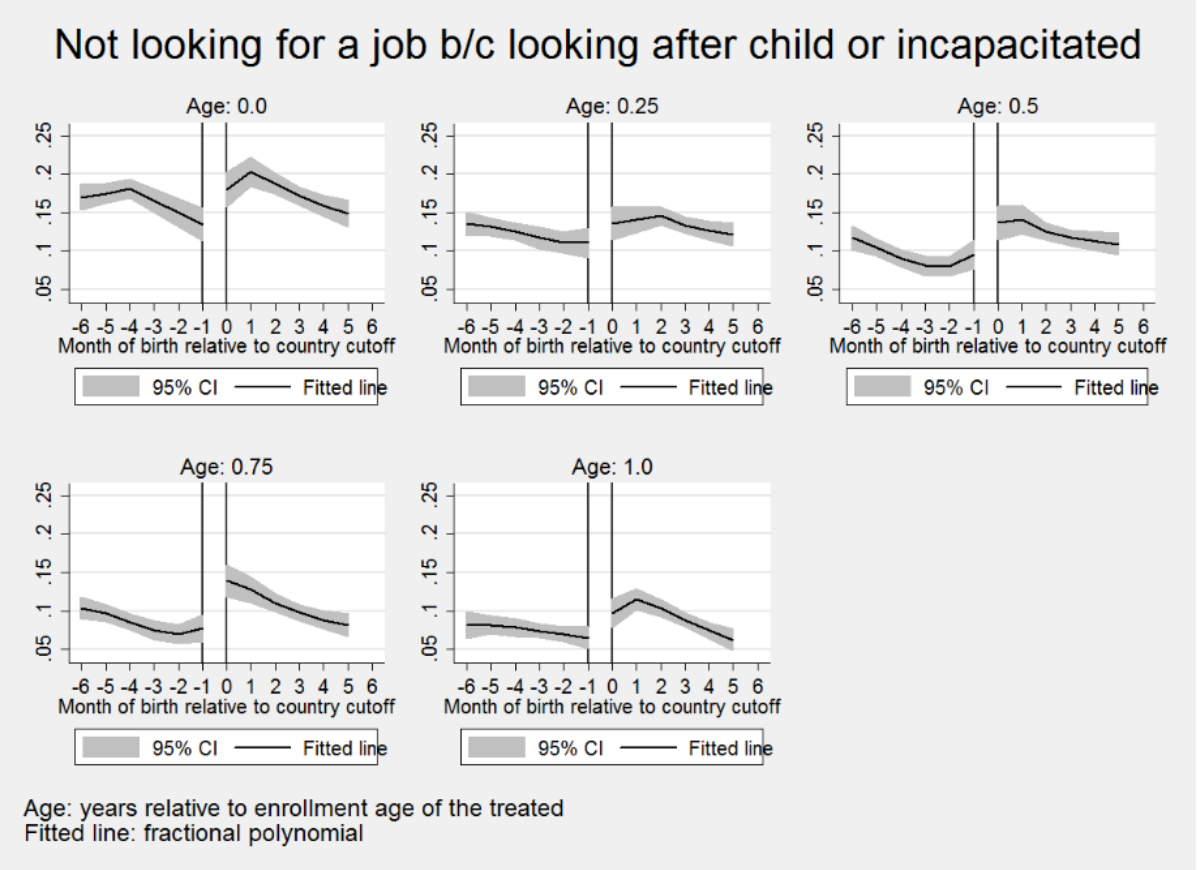
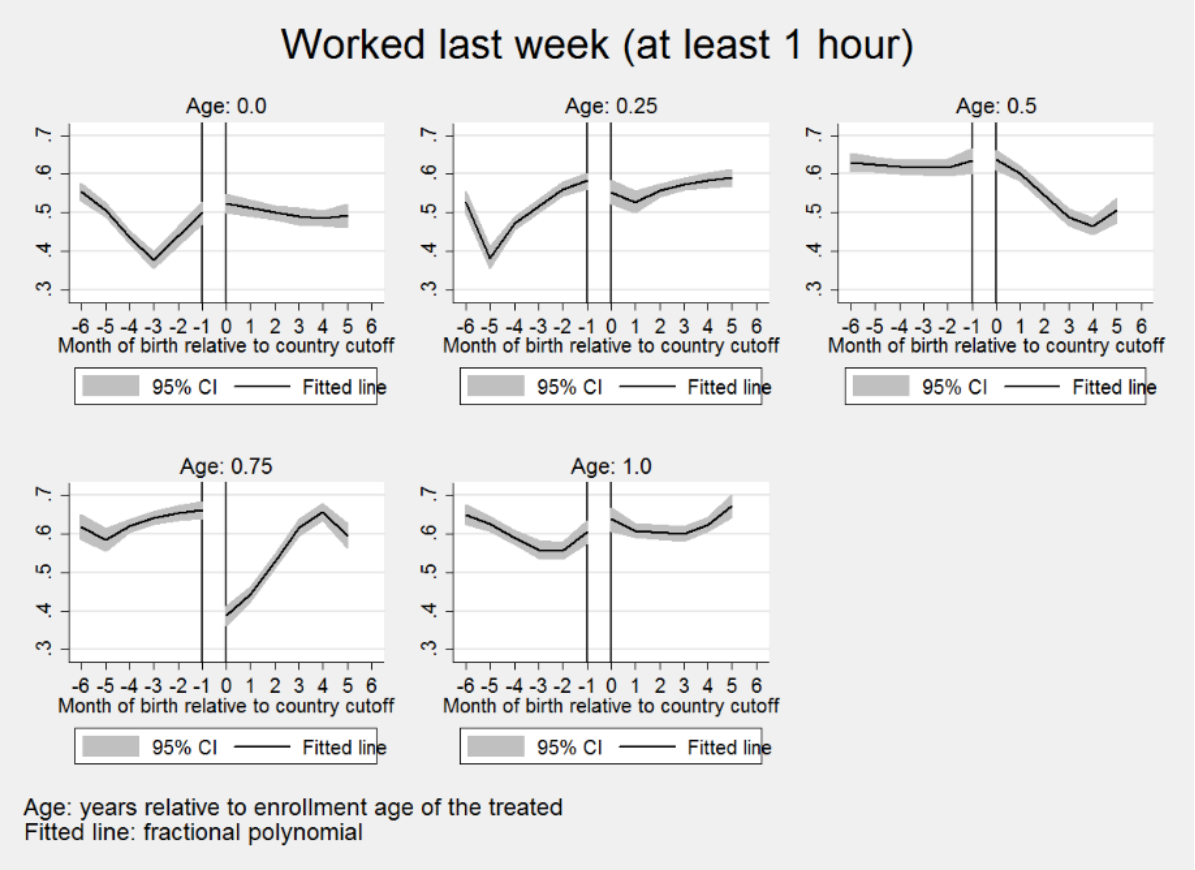


Figure 3: Worked last week (mothers with children born before and after the cutoff date)



Results from the single country regressions (Eq. 2) are reported in Table 4. As indicated in the bottom row, the kindergarten eligibility cutoff points are either on the 1st of January or the 1st of March, depending on the country. The estimated treatment effect is positive and significant for the Czech Republic, France, Hungary and Slovakia. For Austria and Greece the results show insignificant, though still positive effects. These results are in line with what we see on Figure A1, as we see an increase in LFP around age 3 in each country except in Austria and Greece. In order to interpret these results, the bottom of the table shows how much treatment increases childcare availability for the treatment group relative to the control group. We can see large increases in each country, ranging from 44 to 65 percentage points. A rudimentary analysis of the magnitude of the effects is provided in the table, using a Wald estimator of the following form:

$$\beta_W = \frac{E(r_g|T=1)-E(r_g|T=0)}{E(C_g|T=1)-E(C_g|T=0)} \quad (4)$$

We proxy the country-specific childcare availability of the treatment and the control groups with the childcare enrollment rates of 3 and 0-2-year-olds respectively (reflecting country averages of kindergarten and nursery school enrollment rates).

Table 4: Reduced form estimates by country (Eq. 2)

		Austria	Czech Rep.	France	Greece	Hungary	Slovakia
Regression results	*T*	.034	.090	.084	.036	.135	.162
	P-value	.244	.001	.002	.179	.000	.000
	Individual and family controls included	Yes	yes	yes	Yes	yes	yes
	N	771	978	878	1255	1596	629
	R^2	.219	.375	.198	.194	.307	.208
Childcare statistics	*Nursery school enrollment rate*[1]	17%	6%	51%	21%	9%	5%
	Kindergarten enrollment rate[2]	60%	60%	98%	67%	74%	60%
	Difference in childcare availability	44pp	54pp	47pp	45pp	65pp	56pp
	Wald estimate (β_W)	0.079	0.167	0.179	0.078	0.208	0.295
	Birthdate cutoff[3]	March 1	Jan 1	Jan 1	March 1	Jan 1	Jan 1

Notes: [1] [2] OECD Family Database statistics: participation rates for 0-2 and 3 year olds in formal childcare and pre-school services. Data generally include children in center-based services, organized day care and pre-school (both public and private) and those who are cared for by a professional childminder, and exclude informal services provided by relatives, friends or neighbors. Exact definitions may however differ slightly across countries. [3] Cutoff dates based on country expert surveys.

Results of the pooled regression of Eq. (3) are presented in columns (1) and (2) of Table 5. The standard errors, reported in parentheses, are calculated using the Huber-White estimator, and are clustered at the NUTS region level (there are 46 clusters), thus, they are robust to heteroscedasticity.

Model (1) shows how the various elements of the institutional environment affect relative labor supply. The results indicate that it is the length of leave which has the far largest influence on maternal LFP. The shorter the leave, the higher the probability of participating in the labor market around the age of 3 years of the child. In the second specification, we add the interaction terms in order to uncover the interdependencies between childcare and other institutional elements. The estimates suggest that the childcare availability increases maternal labor supply more if the replacement rates of cash benefits are higher. By the same token, in the presence of part-time job opportunities the childcare may have higher significant positive effect on labor supply. Also, if the length of leave is 1-2 years, the childcare effect may be lower. The interactions of the other variables do not appear to have a significant effect.

Table 5: Pooled model specification results (Eqs. 3 and 6)

Dep. var.: LFP	**(1)**	**(2)**	**(3)**	**(4)**
Treatment	0.007	-2.000*	-0.003	-2.009*
	(0.010)	(1.183)	(0.010)	(1.181)
Leave repl. rate	-0.009	-0.028	-0.010	-0.025
	(0.015)	(0.017)	(0.014)	(0.016)
*T*Leave repl. rate*		0.037*		0.036*
		(0.019)		(0.019)
Part-time jobs (%)	0.346	-0.256	0.239	-0.304
	(0.459)	(0.524)	(0.443)	(0.516)
*T*Part-time jobs (%)*		1.220**		1.264**
		(0.585)		(0.587)
Family > Work	-0.722	-0.851	-0.709	-0.858

Dep. var.: LFP	**(1)**	**(2)**	**(3)**	**(4)**
	(0.604)	(0.754)	(0.614)	(0.767)
*T*Family > Work*		0.194		0.201
		(1.313)		(1.298)
Leave: 0-1 yrs[(a)]	0.135**	0.179***	-0.169**	-0.148*
	(0.054)	(0.063)	(0.076)	(0.086)
*T*Leave: 0-1 yrs[(a)]*		-0.101		-0.063
		(0.118)		(0.117)
Leave: 1-2 yrs[(a)]	0.379	1.356	0.153	0.995
	(0.744)	(0.846)	(0.724)	(0.837)
*T*Leave: 1-2 yrs[(a)]*		-1.977**		-1.961**
		(0.953)		(0.952)
Relative LFP			0.751***	0.753***
			(0.111)	(0.125)
Adj. R2	0.161	0.161	0.167	0.166
F	131.488	131.430	130.193	131.967
AIC	4863.924	4868.993	4834.555	4842.379
N	4821.000	4821.000	4821.000	4821.000
Year FE	yes	yes	yes	yes
Demographic controls	yes	yes	yes	yes
Institutional interactions	no	yes	no	yes

Source: EU-LFS dataset, years 2005-2012.

*Note: The full table of the estimated coefficients are presented in Table A2 in the Appendix. Huber-White standard errors are in parenthesis, and are clustered to regions (there are 46 clusters). Cols. (1) and (2) show the results of eq. 3; (3) and (4) show those of eq. 6. Windows are 5 months wide around the cutoff, the relative age of the children included in the sample is .75. Stars indicate significance as: * p<0.1; ** p<0.05; *** p<0.01.*

(a) In comparison to leaves longer than 2 years.

As depicted in Figure A1, the maternal labor market return process is very different across countries, which may be due to several reasons. First, long-term maternal LFP rates (of mothers with a school-aged youngest child) vary a lot across countries. Second, economic and labor market conditions and the timing of economic cycles may also differ which determine employment prospects and thus influence LFP in the given country. In order to account for these differences, we include information on the long-term maternal LFP rate in our estimation. We calculate relative participation rates:

$$r_g = \overline{LFP_g^{(3)}} / \overline{LFP_g^{(8-15)}} \qquad (5)$$

where individual observations are aggregated in groups (g) on the basis of country of residence, treatment status and highest completed education. $\overline{L_g^{(3)}}$ is the mean LFP rate of mothers in group g, who have a youngest child of age 3. $L_g^{(8-15)}$ refers to similar mothers with a youngest child of schooling age (8-15 years). We refer to this latter as the long term LFP rate. r_g is a sort of inverse distance variable, which indicates how far mothers are from their potential (long term) rate of participation at a given child age. The smaller r_g is (the further mothers are from their long-term participation rate), the larger the potential impact of childcare enrollment on participation. In specifications (3) and (4) of Table 5, we include r_{gi} in the estimation equation, which refers to the relative labor supply of group g which individual i belongs to.

$$LFP_{yci} = \beta_1 T_{yci} + r_{gi} + \alpha_y + \beta_2 I_{yc} + \beta_3{}' T_{yci} * I_{yc} + \pi' X_{yci} + \xi_{ycgi} \quad (6)$$

The estimated coefficients of the institutional interactions change only slightly as a result of this modification, which indicates that it is not the macro level labor market differences of the countries that drive the results.

5. Conclusion

This study estimates the effect of childcare availability on maternal labor supply for 6 European countries with different institutional contexts, and utilizes this variation to learn about the interdependencies of childcare and other factors. We first provide comparable, quasi-experimental estimates – based on eligibility cutoffs – from several countries using harmonized data and a unified methodology. We then estimate a pooled model to quantify the relationship of childcare and other institutional factors. The results from the single country estimates as well as the pooled estimates are in line with previous results and theoretical expectations. Single country estimates indicate that a 10 percentage point increase in childcare enrolment rate raises the maternal LFP rate by 0.78-2.95 percentage points depending on country. These estimates are comparable to those of the existing literature. In response to a 10 percentage point increase in childcare availability, Nollenerger and Rodríguez-Planas (2015) found a 2 percentage point increase in maternal LFP rate in Spain, whereas Lovasz and Szabo-Morvai (2015) (Hungary) found an effect of 1.8 in Hungary, and Bauernschuster and Schlotter (2015) an effect of 3.5 in Germany (Bauernschuster and Schlotter, 2015; Lovasz and Szabo-Morvai, 2013; Nollenberger and Rodríguez-Planas, 2015). Estimates from the pooled country model suggest that higher benefit replacement rates, availability of part-time job opportunities and short leave length increase the childcare effect on maternal LFP.

We are still gathering data on further potential countries with eligibility cutoffs, and we are hoping to add more countries to the analysis, which will add further variation to the institutional contexts. We are also still working on the pooled country specification, and plan to make more clear the choice of institutional variables to be included and their functional specification.

References

Angrist, J.D., Krueger, A.B., 1991. Does Compulsory School Attendance Affect Schooling and Earnings? Q. J. Econ. 106, 979–1014. doi:10.2307/2937954

Bauernschuster, S., Schlotter, M., 2015. Public child care and mothers' labor supply—Evidence from two quasi-experiments. J. Public Econ. 123, 1–16. doi:10.1016/j.jpubeco.2014.12.013

Blaskó, Z., 2011. [Three years at home with the child – but not at any cost - Public opinion regarding mothers' employment.] Három évig a gyermek mellett – de nem minden áron. A közvélemény a kisgyermekes anyák munkába állásáról. Demográfia 54, 23–44.

Boca, D.D., Pasqua, S., Pronzato, C., 2009. Motherhood and market work decisions in institutional context: a European perspective. Oxf. Econ. Pap. 61.

Boeckmann, I., Misra, J., Budig, M.J., 2014. Cultural and Institutional Factors Shaping Mothers' Employment and Working Hours in Postindustrial Countries. Soc. Forces sou119. doi:10.1093/sf/sou119

Budig, M.J., Misra, J., Boeckmann, I., 2012. The Motherhood Penalty in Cross-National Perspective: The Importance of Work–Family Policies and Cultural Attitudes. Soc. Polit. Int. Stud. Gend. State Soc. 19, 163–193. doi:10.1093/sp/jxs006

Cascio, E.U., 2009. Maternal Labor Supply and the Introduction of Kindergartens into American Public Schools. J. Hum. Resour. 44, 140–170. doi:10.3368/jhr.44.1.140

Cascio, E.U., Haider, S.J., Nielsen, H.S., 2015. The effectiveness of policies that promote labor force participation of women with children: A collection of national studies. Labour Econ. 36, 64–71. doi:10.1016/j.labeco.2015.08.002

Cipollone, A., Patacchini, E., Vallanti, G., 2014. Female labour market participation in Europe: novel evidence on trends and shaping factors. IZA J. Eur. Labor Stud. 3, 1–40.

Cukrowska-Torzewska, E., 2015. She Cares and He Earns? The Family Gaps in Poland (Working Paper No. 2015–32). Faculty of Economic Sciences, University of Warsaw.

Del Boca, D., 2002. The Effect of Child Care and Part Time Opportunities on Participation and Fertility Decisions in Italy (IZA Discussion Paper No. 427). Institute for the Study of Labor (IZA).

EC, 2013. Barcelona objectives.

EC, 2008. European Commission - PRESS RELEASES - Press release - Childcare services in the EU [WWW Document]. URL http://europa.eu/rapid/press-release_MEMO-08-592_en.htm (accessed 4.20.16).

Fitzpatrick, M.D., 2010. Preschoolers Enrolled and Mothers at Work? The Effects of Universal Prekindergarten. J. Labor Econ. 28, 51–85. doi:10.1086/648666

Givord, P., Marbot, C., 2015. Does the cost of child care affect female labor market participation? An evaluation of a French reform of childcare subsidies. Labour Econ. 36, 99–111.

Lovasz, A., Szabo-Morvai, A., 2013. Does Childcare Matter for Maternal Labor Supply? Pushing the limits of the Regression Discontinuity Framework (Budapest Working Papers on the Labour Market No. 1313). Institute of Economics, Centre for Economic and Regional Studies, Hungarian Academy of Sciences.

Mills, M., Präg, P., Tsang, F., Begall, K., Derbyshire, J., Kohle, L., Miani, C., Hoorens, S., 2014. Use of childcare in the EU Member States and progress towards the Barcelona targets: Short Statistical Report No. 1. Santa Monica, CA: RAND Corporation.

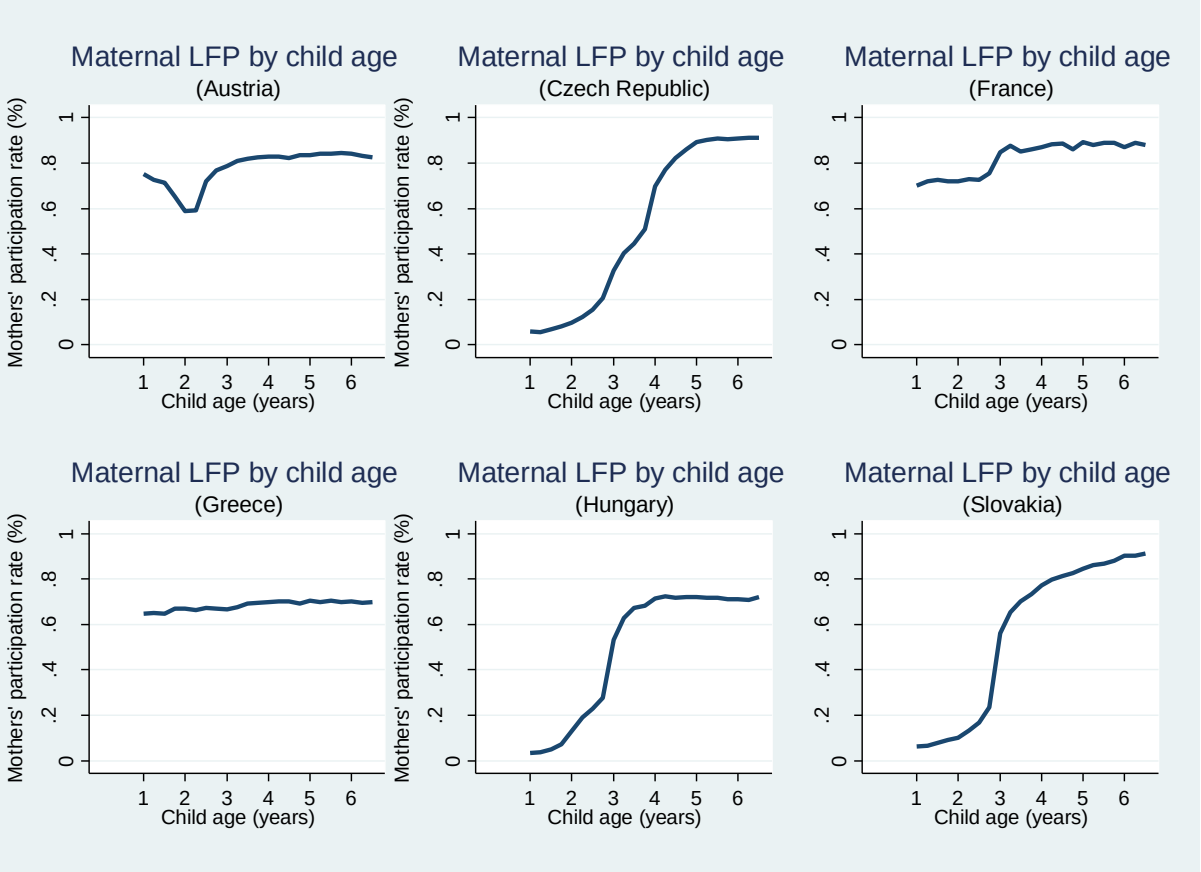
Nollenberger, N., Rodríguez-Planas, N., 2015. Full-time universal childcare in a context of low maternal employment: Quasi-experimental evidence from Spain. Labour Econ. 36, 124–136.

Ray, D., Robson, A., 2016. Certified Random: A New Order for Co-Authorship (Working Paper No. 22602). National Bureau of Economic Research.

Thévenon, O., Solaz, A., 2013. Labour Market Effects of Parental Leave Policies in OECD Countries (OECD Social, Employment and Migration Working Papers). Organisation for Economic Co-operation and Development, Paris.

Appendix

Figure A4: Maternal return to the labor market following childbirth by country



Note: According to country-specific regulations, in some countries mothers receiving maternal benefit are regarded as employed and thus counted as participating in the labor market. This is the reason why in case of Austria the LFP rate is higher for mothers of less than 2-year-olds. This does not affect the estimates of this paper, as the cutoffs are in later years.

Table A1: Country cutoff details and sources of information

Country	Birthdate cutoff	Enrollment date	Age at enrollment (treated)	Expert information on cutoff	EU-SILC information on cutoff
AT	March 1	1 Sept	2.5	yes	yes
CZ	January 1	1 Sept	2.7	yes	yes
FR	January 1	1 Sept	2.7	yes	N/A
GR	March 1	1 Sept	2.5	yes	yes
HU	January 1	1 Sept	2.7	yes	yes
SK	Sept 1	1 Sept	3.0	yes	N/A

Table A2: Pooled model specification full set of results (Eqs. 3 and 6)

	(1)	(2)	(3)	(4)
Treatment	0.007	-2.000*	-0.003	-2.009*
	(0.010)	(1.183)	(0.010)	(1.181)
Educ: low[(a)]	-0.266***	-0.266***	-0.154***	-0.154***
	(0.025)	(0.024)	(0.031)	(0.033)
Educ: medium[(a)]	-0.127***	-0.127***	-0.084***	-0.083***
	(0.019)	(0.019)	(0.019)	(0.020)
Widowed[(b)]	0.089***	0.089***	0.085***	0.085***
	(0.030)	(0.029)	(0.029)	(0.029)
Single[(b)]	0.019	0.020	0.017	0.018
	(0.015)	(0.015)	(0.015)	(0.015)
Age	0.050***	0.050***	0.049***	0.050***
	(0.011)	(0.011)	(0.011)	(0.011)
Age squared	-0.001***	-0.001***	-0.001***	-0.001***
	(0.000)	(0.000)	(0.000)	(0.000)
# of children in HH	-0.056***	-0.056***	-0.056***	-0.056***
	(0.007)	(0.007)	(0.007)	(0.007)
# of inactive in HH	0.019	0.019	0.022	0.022
	(0.025)	(0.026)	(0.025)	(0.026)
Leave repl. rate	-0.009	-0.028	-0.010	-0.025
	(0.015)	(0.017)	(0.014)	(0.016)
*T*Leave repl. rate*		0.037*		0.036*
		(0.019)		(0.019)
Part-time jobs (%)	0.346	-0.256	0.239	-0.304
	(0.459)	(0.524)	(0.443)	(0.516)
*T*Part-time jobs (%)*		1.220**		1.264**
		(0.585)		(0.587)
Family > Work	-0.722	-0.851	-0.709	-0.858
	(0.604)	(0.754)	(0.614)	(0.767)
*T*Family > Work*		0.194		0.201
		(1.313)		(1.298)
Leave: 0-1 yrs[(c)]	0.135**	0.179***	-0.169**	-0.148*
	(0.054)	(0.063)	(0.076)	(0.086)
*T*Leave: 0-1 yrs(c)*		-0.101		-0.063
		(0.118)		(0.117)
Leave: 1-2 yrs[(c)]	0.379	1.356	0.153	0.995
	(0.744)	(0.846)	(0.724)	(0.837)
*T*Leave: 1-2 yrs(c)*		-1.977**		-1.961**
		(0.953)		(0.952)
Relative LFP			0.751***	0.753***
			(0.111)	(0.125)
Constant	0.396	1.399	0.008	0.883
	(0.878)	(0.998)	(0.856)	(0.991)
Adjusted R2	0.161	0.161	0.167	0.166
F	131.488	131.430	130.193	131.967

AIC	4863.924	4868.993	4834.555	4842.379
N	4821.000	4821.000	4821.000	4821.000

Source: EU-LFS dataset, years 2005-2012.

*Note: Huber-White standard errors are in parenthesis, and are clustered to regions (there are 46 clusters). Cols. (1) and (2) show the results of eq. 3; (3) and (4) show those of eq. 6. Windows are 5 months wide around the cutoff, the relative age of the children included in the sample is .75. Stars indicate significance as: * p<0.1; ** p<0.05; *** p<0.01.*

Omitted variables: (a) Education: high level; (b) Marital status: married; (c) Leave: 2-3.5 years