Does Children's Formal Schooling Foster Mothers' Labor Force Participation? Evidence from Peru.

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Abstract

This paper estimates the causal effect of pre-kindergarten enrollment of children in the labor market outcomes of their mothers. I exploit a quasi-experimental setting in Peru, in which children turning 3 years old before March 31st are eligible to be formally enrolled in pre-kindergarten, while those who were born after April 1st have to wait an additional year. This leads to a large discontinuity in the probability of enrollment to pre-kindergarten. Using the child's exact date of birth as an instrument for enrollment, I estimate that the complier mothers who enroll their children in pre-kindergarten at 3 years old are 44 percentage points more likely to work, than those whose children are not eligible for enrollment.

Keywords: pre-kindergarten enrollment, eligibility rule, maternal employment.

JEL Classification: J13, J16, J21

1 Introduction

Worldwide, female participation in labor markets is significantly lower than male participation[1]. Conditional on working, there are still differences in the type of jobs women and men can access and a persistent gender wage gap. Additionally, there is evidence of gender gaps at the household level, both in the roles that partners take within the household and their outcomes in labor markets (Bertrand, Kamenica and Pan, 2015; Dotti Sani, 2015). Social norms,

[1] Aggregate figures from ILO-KILM suggest that female labor force participation was 50 percent in 2014, 26 percentage points below male labor force participation in the same year.

discrimination and specific maternal roles associated with child-rearing (pregnancy, breastfeeding, etc.) are some of the factors explaining the differences both at the labor market and the household levels. Over the past decades, changes in home production technology, divorce laws, and birth-control technology contributed to the increase of female labor force participation and to the changes in intra-household allocations (Leibowitz, 1974; Voena, 2015; Eckstein, Keane and Lifshitz, 2016). However, men are still perceived as the main source of household income. On the other hand, women are expected to be disproportionately in charge of home production and child-rearing, and parenthood exacerbates the still existing gender gaps (Angelov, Johansson and Lindahl, 2016). The presence of small children in the household might then negatively affect women's careers, since mothers usually face most of the labor market costs associated with child-rearing. Several policies aim to reconcile family and work life of mothers (tax credits for lone mothers, maternity leave policies, childcare subsidies), reducing the trade-off between these two spheres of a woman's life and directly changing the opportunity cost of having a child.

This paper explores how the opportunity cost of having a child might also be affected indirectly by policies aiming to affect children's schooling opportunities at early ages. I exploit a quasi-experimental setting in Peru, in which the eligibility rule for kindergarten enrollment was modified. The new policy specified that, starting in 2011, only children born on or before March 31st could be enrolled in 3 year-old pre-kindergarten (*Inicial de 3 años*), while those born after April 1st had to wait an additional year to be eligible for pre-kindergarten enrollment. Using data from the *Encuesta Nacional de Hogares* (ENAHO)[2] in Peru from 2011 and 2012, this paper studies the effects of this policy on maternal labor market outcomes. A unique feature of my dataset is that it contains information on the exact date of birth of all the individuals in the

[2]In English: National Survey of Households.

survey, including children. This makes it possible to implement a Regression Discontinuity design, in which I consider the mothers of children born just after the eligibility cutoff as a counterfactual for mothers that gave birth to their 3 year-old children just before March 31st. My main hypothesis is that the women whose children were born just before the eligibility threshold are more likely to enroll them in the year they reach the age of three. For this group of mothers, pre-kindergarten acts as an implicit childcare subsidy (Berlinski and Galiani, 2007), which might increase their likelihood to work, since it reduces the need to look for alternative child-care arrangements. If this was the mechanism in place, universal free kindergarten might have effects beyond the cognitive development and schooling of children[3]. This has important policy implications because labor market outcomes for women are not usually considered in the design of education policies (and specifically of eligibility thresholds).

The results in these paper show that eligible children in Peru are 30 percentage points more likely to be enrolled in pre-kindergarten at 3 years old than non-eligible children, which is a large increase given the 18 percent enrollment rate of non-eligible children. This effect is somehow larger than previously found in the literature assessing the effect of kindergarten eligibility thresholds on enrollment, which might reflect differences in baseline rates. For example, eligible 3 year-old children in Germany are 27 percent more likely to attend pre-kindergarten (Bauernschuster and Schlotter, 2015), while in Argentina 3 year-old children are only 5 percentage points more likely to attend pre-kindergarten when fulfilling the eligibility criteria (Berlinski, Galiani and McEwan, 2011)[4].

However, while the literature agrees on the fact that eligibility rules affect

[3] A discussion or assessment on which is the optimal eligibility rule for a child cognitive development is beyond the scope of this paper, and by any means I intend to speak in favor or against any particular threshold.

[4] For 4 year old children, Donovan Fitzpatrick (2010) finds a 7.2 percentage points increase in enrollment (14 percent) in certain U.S. states, while Berlinski, Galiani and McEwan (2011) find a 30 percentage points increase in Argentina.

enrollment, the evidence on maternal labor market outcomes is less conclusive. While some studies found positive effects of pre-kindergarten enrollment on female labor supply (Bauernschuster and Schlotter, 2015; Berlinski and Galiani, 2007; Gelbach, 2002), others found no effects (Donovan Fitzpatrick, 2010). In this sense, I aim to provide an additional piece of evidence, in a novel and recent setting. In particular, I find large effects in female labor force participation, even if not in other outcome variables. Moreover, most of the studies found that the effects are driven by the responses of specific subgroups, such as single mothers (Goux and Maurin, 2010), mothers without younger children (Berlinski, Galiani and McEwan, 2011) or women with more than two children (Nollenberger and Rodríguez-Planas, 2013). I propose to additionally explore heterogeneous effects depending on the household structure, since the presence of the extended family in the household might lead to different results. In particular, the presence of grandparents might alleviate the need to look for alternative childcare arrangements for working mothers or act as an extra source of income for mothers of young children (in case they still work)[5].

Additionally, while most of this literature was focused on developed countries, I explore the effects of kindergarten enrollment on maternal labor market outcomes in a Latin American middle income country. This distinction is relevant, since both female labor market outcomes and enrollment rates largely differ between developing and developed countries, which might alter both the results and the mechanisms in place. Moreover, women in developed countries might have broad access to alternative formal childcare arrangements, as compared to women in developing countries. The only exception are Berlinski, Galiani and McEwan (2011) who analyze these effects in Argentina. However, they only have data until 2001, while data in my analysis come from relatively

[5]My results are usually not significant when I study heterogeneous effects, which might respond to lack of statistical power due to relatively small sample sizes.

recent years (2011 and 2012). Again, this distinction is relevant to the extent that female roles in labor markets and social norms have changed in latest years. Moreover, I study the effect of a recent change in policy, while their analysis is based on a rule that has been in place for several years, which might lead to different behavioral responses (i.e.: even if parents do not time the births in response to the eligibility rule, they might modify their decisions in advance as a response to their already known eligibility status).

The empirical strategy used in this paper is similar to that in Berlinski, Galiani and McEwan (2011) and Donovan Fitzpatrick (2010). A potential limitation of this paper compared to Donovan Fitzpatrick (2010) is that I only have access to survey data, which might be an issue given the relatively smaller sample size. However, survey data might have an advantage relative to administrative data if a large share of the working population is not in the formal sector, which might be the case in a developing country such as Peru. Finally, while both Donovan Fitzpatrick (2010) and Berlinski, Galiani and McEwan (2011) consider the effect on maternal employment of enrollment of 4 year-old children, my sample only comprises children who are 3 years old during the period of analysis. This difference is relevant since while 77.7 percent of 4 year-old children were enrolled in pre-kindergarten in Peru in 2009, only 50.2 percent of 3 year-old children were[6]. Therefore, policies affecting enrollment of 3 year-old children might have a larger effect on maternal employment (if they induce more parents to enroll their children). Moreover, mothers of younger children are likely to have spent less time out of work and suffered a lower skill depreciation, which might make them more likely to return to work.

The paper is structured as follows. Section 2 provides background on the Peruvian education system, focusing on the Initial Level (pre-kindergarten and kindergarten) and presents the data. Section 3 discusses the empirical design.

[6]National Institute of Statistics (Peru), based on ENAHO.

Section 4 presents the results for the first stage and section 5 shows the estimates for female labor market outcomes. Section 6 concludes.

2 Context and Data

The academic year in Peru usually starts during the first week of March and ends in the third week of December. The education system consist of three levels: Initial Education (pre-kindergarten and kindergarten), Primary Education (6 grades)[7] and Secondary Education (5 grades). At all levels, education in public institutions is free.

The Initial Education has two cycles: a first cycle, for 0-2 year-old children (which has the role of childcare) and a second cycle for 3 to 5 year-old children, with a total of three levels. For the remainder of the analysis, I will use the terms "Pre-K3", "Pre-K4" and kindergarten to refer to the first, second and third levels of the second cycle of Initial Education, respectively. Even if only kindergarten is mandatory for enrollment in Primary School, enrollment starting at Pre-K3 is strongly encouraged by the Ministry of Education (MINEDU).

The directions for the 2010 academic year[8] established that, starting in 2010, only children reaching the age required for the corresponding level of Initial Education at the beginning of the academic year (March 1st) were eligible for enrollment. Otherwise, parents had to wait one year to enroll their children in kindergarten or pre-kindergarten. For example, a child that was born on February 28th, 2007 was eligible to start Pre-K3 in 2010, while a child born on March 2nd, 2007 had to wait until 2011 to be eligible for enrollment in Pre-K3. This policy faced strong resistance and the eligibility threshold was finally extended to March 31st starting in the 2011 academic year[9]. This study

[7] In order to be eligible to start the first grade of primary education, children have to be 6 years old before June 30th.

[8] See Resolution 0341-2009 of Ministry of Education of Peru.

[9] As I am writing this paper, a new regulation for the academic year 2017 extended the

will cover the period of 2011 and 2012, right after the implementation of the March 31st threshold. Therefore, during the period of analysis, children at Initial Education were eligible for enrollment at the corresponding school level if they were born on or before March 31st, while children born on April 1st and after had to wait an additional year to be eligible for enrollment. The sample of children (and their mothers) turning 3 years old in 2010 will be excluded from this study, since the March 1st cutoff was not actually enforced during 2010, which might affect the validity of the analysis. The eligibility rule established in 2011 was not specific for Pre-K3 but common to all levels of the second cycle of Initial Education. However, I will limit my analysis to 3 year-old children, since this policy is more likely to affect the behavior of these children's parents in the extensive margin (i.e.: parents of non- eligible 4 year-old children might decide to enroll them in Pre-K3, the previous level of formal school to which these children are eligible, while parents of non-eligible 3 year-old children might be more likely to wait one additional year, rather than enroll them in childcare).

Data for this analysis come from the 2011 and 2012 waves of the *Encuesta Nacional de Hogares* (ENAHO), a national-level representative household's survey, conducted by the Peruvian Institute of Statistics starting in 1997. The ENAHO contains rich information on demographic characteristics of the household members, education, health and employment. Moreover, a unique feature of the data is that it includes the exact date of birth of every individual in the sample, including children, which is key for my analysis.

My final sample consists of 1,067 children who turned 3 years old between February and May of either 2011 or 2012[10], and their respective mothers. For this sample, I have information on kindergarten enrollment and on mother's employment status. I have dropped from the sample a small number of obser-

eligibility threshold to July 31st, after several years of an intense debate regarding the adequate cutoff.

[10]These children were born between February and May of 2008 and 2009, respectively.

vations, for which I was not able to link the child to the mother (i.e.: mother was not living in the household), and a small number of inconsistent cases, in which there more than two children linked to the same mother, who were born with less than seven months of difference and were not twins.

A main limitation of the data is that the Education module of the survey was only administered to children who had already reached the age of three when the household was surveyed. Since the survey was conducted along the calendar year, some children turning three between February and May might not be included in the final sample due to the timing of the interview[11].

In the remainder of the paper, I will refer to the subsample of children (and mothers) born in February and March as the "eligible sample" and to the subsample of those born in April and May as the "non-eligible sample".

3 Empirical Strategy

In order to assess the effect of pre-kindergarten enrollment of children on labor market outcomes of mothers, we are interested in estimating the following model for mother i, on region j and survey year t:

$$Y_{itj} = \alpha_0 + \alpha_1 PreK3_i + \boldsymbol{X_i'\delta} + \gamma_t + \lambda_j + \epsilon_{itj} \tag{1}$$

where Y_{itj} measures the labor market outcome of interest (such as full time employment, part time employment, wages, etc), $PreK3_i$ is a dummy variable that takes value 1 if the child is enrolled in pre-kindergarten, and X_i is a vector of children, women and household's characteristics. γ_t captures year fixed effects and λ_j captures region fixed effects. However, estimating this equation by OLS

[11] For example, the Education Module will not be administered to a child turning three in March, if the child belongs to a household that was surveyed in January or February, and therefore, will be excluded from the final sample.

might lead to an inconsistent estimator of the parameter of interest, α_1. Labor market outcomes and kindergarten enrollment might be jointly determined. $E(PreK3_i\epsilon_{itj}) \neq 0$ might also be a concern if kindergarten enrollment is correlated to some omitted variables which also affect female performance in labor markets, such as parental motivation (Berlinski, Galiani and McEwan, 2011).

In order to overcome these concerns and causally identify the effect of pre-kindergarten enrollment on the mother's labor market outcomes, I propose to implement a Regression Discontinuity (RD) design. Similar strategies were used by Fitzpatric Donovan (2010), Berlinski, Galiani and McEwan (2011) and Bauernschuster and Schlotter (2015) to study the effects of pre-kindergarten eligibility rules in different settings.

As discussed before, the Peruvian policy implemented in 2011 establishes a rule based on which the exact date of birth of a child determines her eligibility status for pre-kindergarten enrollment. Under this policy, 3 year-old children born on or before March 31st are eligible for Pre-K3, while children born after April 1st have to wait an additional year. This policy is likely to induce a large discontinuity in the probability of enrollment of children born just before the threshold, compared to children born just after the cutoff. This first stage can be estimated by OLS from the following model:

$$PreK3_{itj} = \beta_0 + \beta_1 E_i + \beta_2 E_i Dist_i + \beta_3 (1 - E_i) Dist_i + \boldsymbol{X_i'\delta} + \gamma_t + \lambda_j + \mu_{itj} \quad (2)$$

where E_i is a dummy variable which takes value 1 for children born before the eligibility cutoff, and 0 otherwise. The parameter β_1 captures the discontinuity in pre-kindergarten enrollment as a consequence of the eligibility rule. $Dist_i$ is a variable indicating the distance of the date of birth of the child from the eligibility cutoff, centered around March 31st[12] (in particular, $Dist_i = 0$ for

[12]This variable is defined as in Berlinki, Galiani and McEwan (2011), but for a different

a child born on March 31st, $Dist_i = -58$ for a child born on February 1st[13] and $Dist_i = 61$ for a child born on May 31st). We expect the estimator of β_1 to be positive in equation (2) since children born before March 31st fulfill the eligibility requirement, and are therefore more likely to be enrolled in formal pre-kindergarten education than those born after April 1st. β_2 and β_3 capture potential differential linear trends at both sides of the eligibility cutoff.

This is a fuzzy RD design since the eligibility rule is not expected to perfectly predict pre-kindergarten enrollment. Parents of children born before March 31st might still choose not to enroll them in pre-kindergarten, even if they are eligible. This might happen if parents consider that their children are too young and prefer to wait an additional year, or if they have younger children in the family and then prefer to take care of both children at home. Parents of children born after April 1st might try to find alternatives to enroll them anyway or enroll them in the first cycle of Initial Education level[14]. The main identification assumption behind the RD strategy is that children are not selectively sorted around the cutoff. This requires that mothers did not time the birth of children in response to the implementation of the eligibility rule. This assumption seems to be plausible in this setting. First, the first attempt to impose a new eligibility threshold was made in 2010, two years after the 2011 eligible children were born, and almost three years after these children were conceived[15]. To the best of my knowledge, no information in the media suggests that the change in the eligibility rule could have been anticipated by these parents, but that this was an unexpected policy for parents that were expecting to enroll their now non-eligible children. Moreover, if mothers were able to time the births, we would

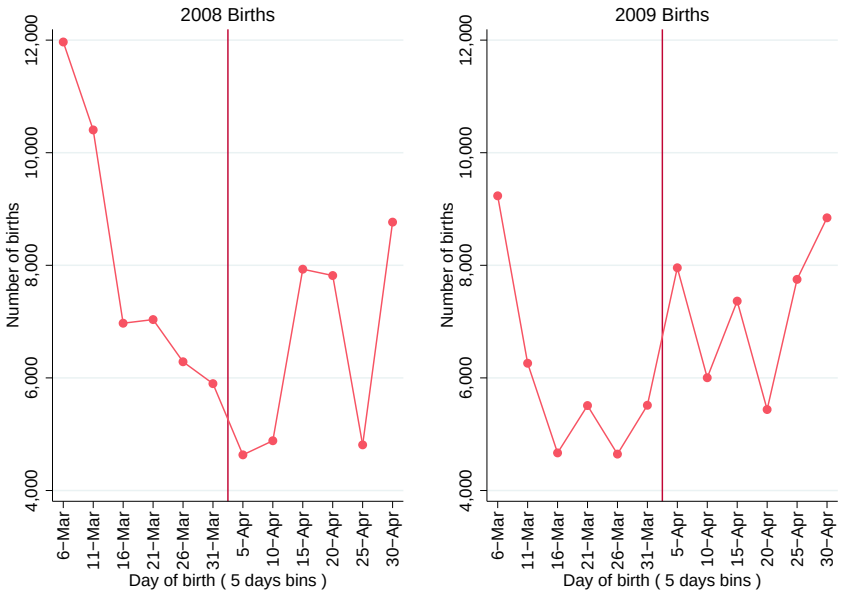
eligibility rule.

[13] For children born in 2008, $Dist_i = -59$ at February 1st, since 2008 was a leap year.

[14] This might not be a problem in our data, since the first cycle of Initial Education is not considered within the Education module of the ENAHO. Therefore, these children should be reported as not enrolled.

[15] For the eligible sample of 2012, the change in policy occurred at least one year after the children were born and about two years after conception.

Figure 1: Number of births in March and April 2008 and 2009 (5 days bins)



Source: ENAHO 2011 and 2012.
Notes: Number of births in March and April of 2008 and 2009, grouped in 5 days bins (i.e.: data point for March 31st groups all births from March, 26th to March, 31st). First bin of March groups births for the first 6 days of the month. Population weights coming from the ENAHO used to compute the number of births.

expect an spike in births at the end of March as compared to the beginning of April, which is not supported by data coming from the ENAHO[16]. Figure 1 shows the frequency of births during March and April of 2011 and 2012, grouped in 5 days bins. As can be seen from this figure, there is no obvious jump in the number of births between the end of March and the beginning of April, neither in 2011 nor in 2012[17].

Data from the Ministry of Health at the month and province level support

[16] For example, in my sample, there were 43 births in the last ten days of March and 37 during the first ten days of April 2008. In 2009, however, only 33 babies in the sample were born during the same period of March, but 50 during the first 10 days of April.

[17] I could also expect an spike in the numbers of births at the end of March if, for example, April 1st were a weekend. March 31st was a Monday in 2008, and a Tuesday in 2009, which might alleviate this concern.

the hypothesis of no spike of births in March as compared to April (see Figure A1 in the Appendix). These pieces of evidence, even if not ideal, support the assumption of random sorting around the cutoff[18].

Identification may also fail if parents of non-eligible children were more likely to misreport their date of birth in order to enroll them. Once again, it is not likely that these parents could have anticipated the policy and therefore, that they misreported the date of birth of their children when they were born. It would be more likely that they attempt to misreport the dates of birth of the child to school authorities. However, the child's birth certificate or citizenship ID must be presented at school upon enrollment, which might alleviate this concern. Moreover, even if this was the case, this would not be a concern if the child's exact date of birth is not misreported in the survey (these children will be part of the "always takers"subsample, but their existence does not threaten the identification). Finally, identification requires that covariates affecting labor market outcomes other than kindergarten enrollment, behave smoothly around the cutoff. In this case, individuals at the right of the cutoff can be used as a counterfactual of individuals just at the left of the cutoff. Table 1 shows descriptive statistics for the eligible subsample (children born in February and March, and their mothers) and the non-eligible subsample (children born in April and May, and their mothers). There are not statistically significant differences in any of the relevant variables across subsamples.

[18] Ideally, I would like to have the daily number of births at the national and regional level coming from administrative registers, but I do not have this data available at this moment.

Table 1: Descriptive Statistics Eligible vs. Non-Eligible

Variable	Eligible	Non-eligible	Difference	p-value
Child is a boy	0.517	0.501	-0.015	0.619
	(0.500)	(0.500)		
Mother Age	31.166	30.564	-0.602	0.198
	(7.602)	(7.696)		
Indigenous	0.748	0.790	0.042	0.100
	(0.435)	(0.407)		
No High School Degree	0.560	0.568	0.007	0.807
	(0.497)	(0.496)		
High School Degree	0.227	0.233	0.006	0.813
	(0.419)	(0.423)		
Tertiary Education	0.128	0.121	-0.007	0.712
	(0.335)	(0.327)		
Some College	0.028	0.037	0.008	0.442
	(0.166)	(0.188)		
College and +	0.055	0.040	-0.014	0.266
	(0.228)	(0.197)		
Married	0.820	0.798	-0.022	0.354
	(0.384)	(0.384)		
Coastal region	0.259	0.244	-0.015	0.572
	(0.439)	(0.430)		
Hills region	0.381	0.366	-0.015	0.615
	(0.486)	(0.482)		
Jungle region	0.256	0.274	0.018	0.500
	(0.437)	(0.446)		
Metropolitan Lima	0.104	0.115	0.011	0.543
	(0.306)	(0.320)		
HH size	6.048	6.158	0.111	0.431
	(2.284)	(2.322)		
Number of HH members, 18+	3.030	3.156	0.126	0.174
	(1.496)	(1.535)		
Number of HH members, 65+	0.284	0.276	-0.008	0.854
	(0.774)	(0.674)		
Number of HH members, 0-5	0.473	0.509	0.036	0.421
	(0.712)	(0.738)		
Number of HH members, 6-17	1.524	1.467	-0.058	0.480
	(1.322)	(1.349)		
Observations	526	542		

Source: ENAHO 2011 and 2012.

Notes: Final sample includes all children and their mothers for which the date of birth of the child is within two months of the cutoff (children born between February and May). Summary statistics pool observations from the survey waves of 2011 and 2012. Final sample includes all observations with non-missing values in employment. A full description of the listed variables can be found in Appendix B. Standard errors in parenthesis.

To the extent that the previous assumptions are satisfied, I can use the exact date of birth of a child as an instrument for pre-kindergarten enrollment, to estimate equation (1) by TSLS. For the instrument to be valid, the two usual instrumental variables conditions must hold. First, the instrument might be relevant to explain the enrollment in pre-K3, which I will discuss in next section. Second, the exclusion restriction has to hold, requiring that $E(E_i\epsilon_{itj}) = 0$. Under this assumption, the date of birth of a child should not affect maternal labor market outcomes directly, but only through its effect on pre-kindergarten enrollment. In this setting, the IV estimator of $\alpha_1$does not recover the Average Treatment Effect, but the Local Average Treatment Effect (LATE) (Angrist, Imbens and Rubin, 1996). This parameter measures the average treatment effect on the *compliers*, whose behavior was induced by the policy[19]. These results cannot be generalized to the population of *non-compliers* (*never takers* and *always takers*), who do not change their behavior in response to the instrument (Galiani, Rossi and Scahrgodsky, 2011).

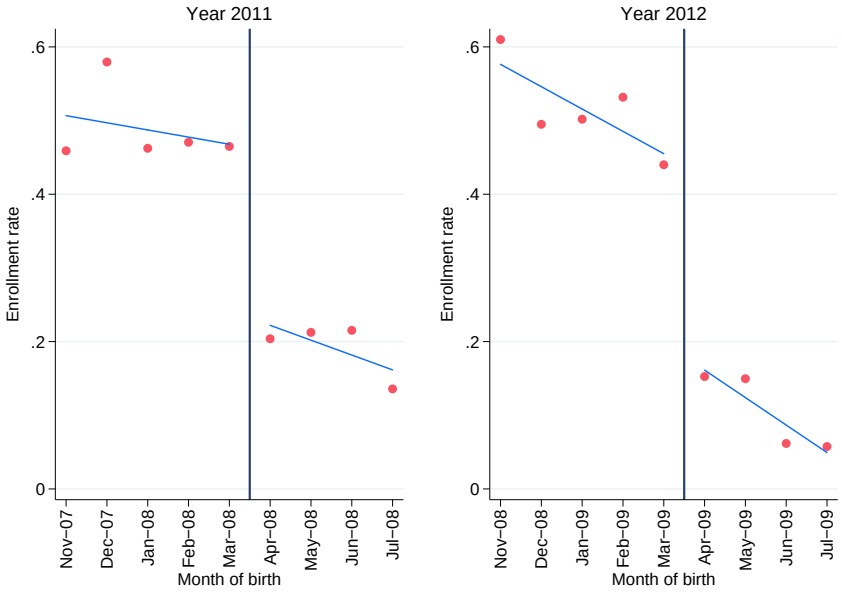
4 Pre-kindergarden Enrollment

This section explores how the date of birth of a child affects her probability of being enrolled in pre-kindergarten.

Figure 2 shows a clear discontinuity in Pre-K3 enrollment around the eligibility threshold both in 2011 and 2012. While 46 percent of the 3 year-old eligible children born in March are enrolled in Pre-K3 in 2011, only 20 percent of the non-eligible children born in April are. In 2012, 44 percent of the eligible

[19]The *compliers* are the subsample of individuals who enroll their children in kindergarten only because they were eligible, but would not have enrolled them in pre-kindergarten otherwise, and those who did not enroll the children because they were not eligible but would have enrolled the children in pre-kindergarten if they would have met the requirements.

Figure 2: Enrollment rates by date of birth (2011 and 2012)



Source: ENAHO 2011 and 2012
Notes: Each figure plots the enrollment rate in Pre-K3 for 2011 and 2012, respectively, as a function of the month of birth of the children. The vertical line stands for the eligibility cutoff, set in March 31st.

children are enrolled, while only 15 percent of the non-eligible children are. As discussed before, it is possible that non-eligible children were actually enrolled in the previous pre-kindergarten level (for 2 year-old children). This does not seem to be the case, since the overwhelming majority of enrolled children in my sample are reported to be in grade 1 of Initial Education (which is associated to Pre-K3). Moreover, as mentioned above, children in the first cycle of Initial Education should not be reported as enrolled in formal education in the Education module of the ENAHO. It is possible that parents misreport the actual date of birth upon enrollment to make children eligible. This does not seem plausible, since as discussed above, parents must submit the birth certificate of the child to the school authorities, which states the exact date of birth. The most plausible explanation is that, in some cases, the eligibility threshold was not strictly enforced. There is not additional data on the survey which allows disentangling these potential explanations for enrollment of non-eligible children. However, this should not affect my TSLS estimators[20]. Table 2 shows the OLS first stage estimators of equation (2), pooling the 2011 and 2012 observations. All specifications consider a bandwidth of two months around the eligibility threshold (i.e.: mothers of children born in April and May are used as a control group for mothers of children born on February and March), placing equal weight to all observation within this bandwidth. Estimation in column 1 only includes the trend terms at both sides of the eligibility threshold. The estimator of $\beta_1$is large, positive and statistically significant at the 1 percent level, suggesting that eligible children are 30 percentage points more likely to be enrolled in Pre-K3 than non-eligible children. Results are robust to controlling for the gender of the children (column 2) and for the characteristics of the mother, such as age, educational level, whether the mother is indigenous and whether the mother

[20]These children and their mothers will be *always takers*,since the instrument does not induce a change in their behavior.

is married or cohabiting (column 3). Column 4 additionally controls for the household's structure. In my preferred specification, presented in column 5, I introduce survey year and regional fixed effects, which do not alter the point estimation and significance level of β_1.

Table 2: First Stage. Pre-Kindergarten enrollment by eligibility status.

	(1)	(2)	(3)	(4)	(5)
Eligible	0.302***	0.304***	0.284***	0.290***	0.292***
	(0.079)	(0.079)	(0.073)	(0.074)	(0.073)
Children Char.	No	Yes	Yes	Yes	Yes
Mother Char.	No	No	Yes	Yes	Yes
HH structure	No	No	No	Yes	Yes
Region & year FE	No	No	No	No	Yes
Observations	1,067	1,067	1,067	1,067	1,067

Source: ENAHO 2011 and 2012.
Notes: OLS regressions of kindergarten enrollment on a dummy variable indicating whether the child was born before the eligibility threshold. All regressions include linear trend terms at both sides of the threshold, as in Equation (2). The sample includes children born within two months from the threshold. Survey weights are used in all regressions. Heteroscedastic-robust standard errors in parenthesis.
* Significant at 10%, ** Significant at 5%, *** Significant at 1%

Since the academic year starts in March, households surveyed in January and February are more likely to report that children are not enrolled even when parents are planning to enroll them during the same calendar year. Column 1 of Table 3 shows that the estimator of β_1 slightly increases in magnitude (0.33) when I restrict the sample only to households surveyed during the academic year (March - December). The remaining columns of Table 3 replicate the estimation in column 5 of Table 2 for different bandwidths around the eligibility threshold. In column 2, I extend the bandwidth from 2 to 3 months (including the children born in January and June), which does not affect the significance or magnitude of the coefficient compared to my preferred specification. In column 3, I only include children born within a month of the cutoff (those born in March and April). The coefficient increases to 0.31 and remains significant at 1 percent. In

the last two columns of Table 3, I perform two different falsification tests. In column 4, I assume the eligibility cutoff to be May 31st, and I use the children born in June and July as a control group for those born in April and May. As expected, this placebo threshold does not predict enrollment. In column 5, I assume the threshold to be at the end of January, considering children born either in February or March as the control group for those born in December and January. In this case, I expect the coefficient to be positive, since the younger the child the more likely the parents will decide to wait one additional year for enrollment, even if they can formally do so. As expected, the coefficient is positive but small in magnitude and not significant.

Table 3: Pre-Kindergarten enrollment by eligibility status. Different samples.

	(1)	(2)	(3)	(4)	(5)
Eligible	0.328***	0.280***	0.307***	-0.044	0.056
	(0.081)	(0.058)	(0.087)	(0.122)	(0.111)
Full set controls	Yes	Yes	Yes	Yes	Yes
Observations	892	1,606	543	1,053	1,106

Source: ENAHO 2011 and 2012.
Notes: OLS regressions of kindergarten enrollment on a dummy variable indicating whether the child was born before the eligibility threshold, based on Equation (2). All regressions include linear trend terms to the left and the right of the threshold (see Equation (2)). Sample in column 1 does not include children in households which were surveyed during January and February (see text). Sample in column 2 includes children born within three months from the cutoff. Sample in column 3 only include children born in March and April. Column 4 considers a placebo threshold in May 31st and column 5 considers a placebo threshold in January 31st. In both cases, samples include children within two months from the threshold. All regressions include the same full set of controls as specification in column 5 of Table 2. A full list and description of these control variables can be found in Appendix B. Survey weights are used in all regressions. Heteroscedastic-robust standard errors in parenthesis.
* Significant at 10%, ** Significant at 5%, *** Significant at 1%

To sum up, the eligibility rule based on the exact date of birth of the children is a key factor explaining enrollment to pre-kindergarten. Even if this level of

Initial Education is not mandatory, there is a large discontinuity in enrollment rates around the eligibility threshold, with parents willing to enroll their children even at the young age of 3 years old, provided they can do so. Therefore, those children who meet the age requirement implemented in 2011 are about 30 percentage points more likely to be enrolled than those who do not meet the eligibility criteria.

5 Labor Market Outcomes

This section presents the estimators of the effect of pre-kindergarten enrollment of 3 year-old children on the labor market outcomes of their mothers.

I start by presenting the reduced form estimators of the mother's labor market outcomes (Y_i) on a dummy indicating whether her child is eligible for Pre-K3 (E_i), corresponding to ρ_1 in the following model:

$$Y_{itj} = \rho_0 + \rho_1 E_i + \rho_2 * E_i * Dist_i + \rho_3 * (1 - E_i) * Dist_i + \boldsymbol{X_i'\delta} + \gamma_t + \lambda_j + \nu_{itj} \quad (3)$$

Column 1 in Table 4 shows the results of estimating equation (3) with no additional controls. Mothers of eligible children are 13 percentage points more likely to be employed, and the estimator is significant at 10 percent. This result is robust to including the full set of controls used in column 5 of Table 2 (column 2). Results are also robust to restricting the sample to households surveyed between March and December (column 3). However, when I extend the sample to include the mothers of children born within 3 months of the eligibility threshold (column 4), the coefficient decreases in magnitude and becomes not significant (p-value=0.103). Finally, the reduced form estimator increases in magnitude (to 0.20) and becomes significant at 5 percent when the sample is restricted to

the mothers of children born within a month of the cutoff (column 5).

Table 4: Mother's labor force participation by child's eligibility status.

	(1)	(2)	(3)	(4)	(5)
Eligible	0.122	0.129*	0.131*	0.096	0.207**
	(0.079)	(0.072)	(0.075)	(0.059)	(0.092)
Full set controls	No	Yes	Yes	Yes	Yes
Observations	1,067	1,067	892	1,606	543

Source: ENAHO 2011 and 2012.
Notes: OLS regressions of mother's labor force participation on a dummy indicating whether the child was born before the formal enrollment eligibility threshold, based on Equation (3). All regressions include linear trend terms to the left and the right of the threshold (see Equation (3)). Specification in column 1 does not include additional controls. Specifications in column 2-5 include the same full set of controls as in specification 5 in Table 2. A full list and description of control variables can be found in Appendix B. Sample in column 3 excludes households surveyed in January and February. Sample in column 4 includes all children born within three months from the cutoff. Sample in column 5 only include children born in March and April. Survey weights are used in all regressions. Heteroscedastic-robust standard errors in parenthesis.
* Significant at 10%, ** Significant at 5%, *** Significant at 1%

Table 5 shows the TSLS estimators of Equation (1), using the eligibility threshold as an instrument for Pre-K3 enrollment and including separate linear trends at both sides of the cutoff. These estimators are just the reduced form estimators presented in Table 4 divided by the first stage estimators reported in Table 2 ($\alpha_1^{TSLS} = \frac{\rho_1}{\beta_1}$), and measure the causal effect of pre-kindergarten enrollment on maternal labor market outcomes. Column 1 shows the results when no additional controls are included. The coefficient for Pre-K3 enrollment is positive and large (0.40), but not significant at standard levels. Column 2 shows the results including the same full set of controls as in column 5 of Table 2. In this case, the TSLS estimator is positive and large, suggesting that those compliers mothers who enroll their children in pre-kindergarten are 44 percentage points more likely to work. This estimator is significant at 10

percent level[21].

Columns 3-5 of Table 5 show the TSLS results for different samples, as in columns 3-5 of Table 4. The point estimator remains virtually unchanged when I exclude from the sample the households that were surveyed in January and February (column 3). In column 4, I extend the bandwidth to 3 months around the cutoff. The TSLS decreases to 0.34 and becomes not significant (p-value=0.103) . Finally, when we restrict the sample to observations which are only within a month of the cutoff, the coefficient becomes larger in magnitude (0.67) and significant at 5 percent level. Results are robust to using pre-kindergarten attendance instead of pre-kindergarten enrollment in the regressions (not reported).

[21] Results are similar when I only include a partial set of controls (as in columns 3-4 of Table 2), but for the sake of space, I do not report these regressions here. The TSLS estimator becomes not significant when I only include the gender of the child as a control in the regressions.

Table 5: TSLS estimators. Mother's labor force participation and child's enrollment.

	(1)	(2)	(3)	(4)	(5)
Pre-K3 Enrollment	0.403	0.444*	0.401*	0.344	0.674**
	(0.280)	(0.255)	(0.240)	(0.211)	(0.339)
Full set controls	No	Yes	Yes	Yes	Yes
Observations	1,067	1,067	892	1,606	543

Source: ENAHO 2011 and 2012.
Notes: TSLS regressions mother's labor force participation on a dummy indicating whether the child was enrolled in Pre-K3 , based on Equation (1) and including linear trend terms to the left and the right of the threshold. The excluded instrument is a dummy variable indicating whether the child was born before the formal eligibility threshold (March 31st). Specification in column 1 does not include additional controls. Specifications in column 2-5 include the same full set of controls as in specification 5 in Table 2. A full description of these variables can be found in the Appendix B. Sample in column 3 excludes households surveyed in January and February. Sample in column 4 includes children born within three months of the cutoff. Sample in column 5 only include children born in March and April. Survey weights are used in all regressions. Heteroscedastic-robust standard errors in parenthesis.
* Significant at 10%, ** Significant at 5%, *** Significant at 1%

These results suggest that pre-kindergarten enrollment of children has a causal impact on maternal employment. Those mothers who are induced by the eligibility rule to enroll their children in pre-kindergarten at 3 years old are more likely to work than those whose children are non-eligible. However, the results seems to imply that, if children are non-eligible for enrollment in pre-kindergarten, mothers do not seek for alternative childcare arrangements and are more likely to remain out of work.

Table 6 reports the TSLS for different outcome variables. Pre-kindergarten enrollment of eligible children has a positive effect of wages, hours worked, the probability of having a secondary job, and the probability of working full time. However, the coefficients are imprecisely estimated in all cases.

Table 6: TSLS estimators. Mother's labor market outcomes and child's enrollment.

	Total Wage	Sec. Employment	Hours	Works FT
Pre-K3 Enrollment	319.431	0.379	19.289	0.178
	(215.048)	(0.326)	(17.848)	(0.389)
Full set controls	Yes	Yes	Yes	Yes
Observations	527	815	815	815

Source: ENAHO 2011 and 2012.
Notes: TSLS regressions of mother's labor market outcomes on a dummy indicating whether the child was enrolled in Pre-K3, based on Equation (1) and including linear trend terms to the left and the right of the threshold. The excluded instrument is a dummy variable indicating whether the child was born before the formal eligibility threshold (March 31st). All regressions include the same full set of controls as in specification 5 in Table 2. A full description of these variables can be found in the Appendix B. Dependent variable in specification in column 1 measures the total wage of the child's mother. Dependent variable in column 2 is a dummy variable indicating whether the mother has a secondary occupation. Dependent variable in column 3 is total number hours worked in primary occupation. Dependent variable in column 4 is a dummy variable indicating whether the child's mother worked more than 20 hours during the previous week. Survey weights are used in all regressions. Heteroscedastic-robust standard errors in parenthesis.
* Significant at 10%, ** Significant at 5%, *** Significant at 1%

5.1 Heterogeneous effects

I finally discuss heterogeneous effects for different subsamples, and which are the potential mechanisms explaining the differences. Panel (a) of Table 7 shows the first stage estimators for different subsamples, while panel (b) reports the TSLS estimators.

First, enrollment's response to the eligibility rule might depend on whether children have younger (non-eligible) siblings. As expected, 3 year-old eligible children are less likely to be enrolled when there are younger children in the family (column 1 of panel A), compared to the case in which they are the youngest (column 2). Having a younger (non-eligible) child makes the mother more likely to stay at home taking care of this child, which lowers the opportunity cost of not enrolling the eligible child. It might be even costlier for the mother to make

her child to attend pre-kindergarten than to keep the child at home. However, the effects on female labor force participation are positive but not precisely estimated for any of the subsamples (column 1 and 2 of panel B), which might be the consequence of small sample sizes. The eligibility rule only has a significant effect on the enrollment decisions when the child's grandparents are not in the household (column 3 of Panel A). When the 3 year-old child is the grandchild of the head of household, the first stage estimator is small in magnitude and not significant at standard levels (column 4 of Panel A). This might reflect a higher ability to substitute formal childcare with informal childcare arrangements for women living in extended households (i.e.: grandparents might take care of the child while the mother is working). Again, the TSLS estimators are imprecisely estimated for both subsamples (columns 3 and 4 of Panel B). The discontinuous enrollment's response to the eligibility rule seems to be driven by the sample of children whose mothers are married or cohabiting (column 5 of Panel A). For the subsample of non-married women, these effects are small, negative and not significant[22] (column 6 of Table 7). In results not reported, I have also checked for heterogeneous effects in different regions of the country. The eligibility rule plays a larger role determining enrollment in Metropolitan Lima than in the rest of the country. The TSLS estimators are also larger, but again imprecisely estimated.

[22] This might reflect lack of power due to small sample size.

Table 7: Heterogeneous effects.

	(1)	(2)	(3)	(4)	(5)	(6)
	Younger sibling	No younger sibling	No Extended	Extended Family	Partner in HH	No Partner
	Panel (a). First Stage: Children enrollment					
Eligible	0.168	0.320***	0.361***	0.104	0.365***	-0.029
	(0.139)	(0.084)	(0.089)	(0.123)	(0.080)	(0.147)
	Panel (b): TSLS. Mother's labor force participation					
Pre-K3 Enrollment	1.796	0.265	0.116	3.284	0.405*	-0.845
	(1.603)	(0.251)	(0.246)	(3.567)	(0.234)	(6.067)
Full set controls	Yes	Yes	Yes	Yes	Yes	Yes
Observations	203	854	757	310	864	203

Source: ENAHO 2011 and 2012.
Notes: Panel (a) reports the coefficient of OLS regressions of kindergarten enrollment on a dummy variable indicating whether the child was born before the eligibility threshold, based on Equation (2), for different subsamples. Panel (b) reports the coefficient of TSLS regressions of mother's labor force participation on a dummy indicating whether the child was enrolled in Pre-K3 , based on Equation (1), for different subsamples. The excluded instrument is a dummy variable indicating whether the child was born before the formal eligibility threshold (March 31st). All regressions include children within two months from the eligibility cutoff. All regressions include linear trend terms to the left and the right of the threshold and the same full set of controls as in specification 5 of Table 2. A full description of these variables can be found in the Appendix. Columns 1 and 2 split the sample according to whether there is a younger sibling in the family or not. Columns 3 and 4 split the sample according to whether the child's gradparents are present in the HH. Column 5 and 6 divide the sample taking into account the marital status of the child's mother (married or cohabiting vs. single or divorced). Survey weights are used in all regressions. Heteroscedastic-robust standard errors in parenthesis.
* Significant at 10%, ** Significant at 5%, *** Significant at 1%

6 Conclusion

This paper adds evidence on whether having a child enrolled in pre-kindergarten boosts maternal labor market participation. I exploit a quasi-experimental setting in Peru, in which starting in 2011, the Ministry of Education modified the eligibility threshold for pre-kindergarten enrollment. This policy made 3 year-old children born before March 31st eligible for pre-kindergarten, while those who were born after March 31st had to wait an additional year to be formally eligible. This rule led to a large discontinuity in enrollment rates, with eligible

children being 29 percentage points more likely to be enrolled than non-eligible ones. My TSLS estimators suggests that those compliers mothers who enroll their children in pre-kindergarten at 3 years old are 44 percentage points more likely to work than those mothers whose children were not eligible for enrollment.

In a similar setting, Berlinski, Galiani and McEwan (2011) do not find a large impact of the eligibility rule on enrollment of 3 year-old children (or in labor market outcomes for the mothers of these children)[23]. The differences in results might respond to differences in the eligibility cutoff between the two countries studied. While in both cases the academic year starts in March, the eligibility threshold in Argentina is June 30th, while in Peru was March 31st during the period of analysis. At the beginning of the school year, Peruvian mothers face the decision of whether to enroll their 3 year-old children (or those who are close to turn 3 years old), while Argentinian mothers have to decide whether to enroll their 2 years and 8 month-old children. At this young age of rapid change in cognitive abilities, a four months difference can strongly impact the decision regarding the child's enrollment in preschool. Therefore, the exact eligibility cutoff might actually determine the potential of this policy to influence maternal employment. So, these results add to the debate on the optimal eligibility threshold, by showing that this rule can impact the child's family as a whole. Moreover, a comprehensive policy providing non-eligible children with access to childcare institutions and giving parents the adequate incentives to enroll them might also improve maternal employment [24]As I am writing this paper, the Peruvian cutoff for enrollment in pre-kindergarten was modified, and set at

[23]Their effects are concentrated among 4 year-old children, who I do not examine in this paper. I plan to do this in future work.

[24]Currently, mothers can enroll children in the first level of Initial Education. However, it is not clear whether this educational level is broadly offered in public institutions and whether most parents have access to it. Moreover, if parents do not perceive this level as formal education might postpone enrollment until the child is eligible for 3 years old pre-kindergarten.

July 31st. The previous discussion suggests that this change in the eligibility rule might further promote maternal employment, but this potential has limits as the cutoff is moved further away from the previous one[25]. However, it is still too soon to evaluate the impact of this new policy both in enrollment and maternal employment. In future research, I will include the new data to better interpret the effects of eligibility rules.

As a final remark, results in this paper only provide evidence on the link between children's preschool enrollment and *maternal* employment. I have not seen any paper so far that examines the effect of the child's pre-kindergarten enrollment on the labor market outcomes of the *father*. This might be relevant to evaluate, as part of a broad effort to understand how roles are determined within the households in developing countries, and how child-related policies have the potential to redefine them.

7 References

Angelov, N., Johansson, P., and Lindahl, E. (2016). "Parenthood and the Gender Gap in Pay", *Journal of Labor Economics*, 34(3): 545-578.

Angrist, J., Imbens, G. and Rubin, D. (1996). "Identification of Causal Effects Using Instrumental Variables", *Journal of the American Statistical Association*, 91(434): 444–455.

Bauernschuster, S. and Schlotter, M. (2015). "Public child care and mothers' labor supply: Evidence from two quasi-experiments", *Journal of Public Economics*, 123: 1-16.

Berlinski, S. and Galiani, S. (2007). "The effect of a large expansion of pre-primary school facilities on preschool attendance and maternal employment".

[25] As previously mentioned, this paper is silent about which is the optimal threshold in terms of cognitive abilities of the child. My conclusions are only related to the potential impact that eligibility rule might have on maternal labor market outcomes.

Labour Economics, 14: 665-680.

Berlinski, S. Galiani, S. and McEwan, P. (2011). "Preschool and Maternal Labor Market Outcomes: Evidence from a Regression Discontinuity Design.", *Economic Development and Cultural Change*, 59(2): 313-344.

Bertrand, M., Kamenica, E. and Pan, J. (2015). "Gender Identity and Relative Income within Households", *The Quarterly Journal of Economics*, 130(2): 571-614.

Donovan Fitzpatrick, M. (2010). "Preschoolers Enrolled and Mothers at Work? The Effects of Universal Prekindergarten", *Journal of Labor Economics*, 28(1): 51-85.

Dotti Sani, G.M.(2015). "Within Couple Inequality in Earnings and the Relative Motherhood Penalty. A Cross-National Study of European Countries", *European Sociological Review*, 31(6): 667-682.

Eckstein, Z., Keane, M., and Lifshitz, O. (2016). "Sources of Change in the Life-Cycle Decisions of American Men and Women: 1962-2014", Working Paper.

Galiani, S., Rossi, M. and Schargrodsky, E. (2011). "Conscription and Crime: Evidence from the Argentine Draft Lottery", *American Economic Journal: Applied Economics*, 3: 119-136.

Gelbach, J. (2002). "Public Schooling for Young Children and Maternal Labor Supply", *The American Economic Review*, 92(1): 307-322.

Goux, D. and Maurin, E. (2010). "Public school availability for two year-old and mothers' labour supply", *Labour Economics*, 17: 951-962.

Leibowitz, A. (1974). "Home Investments in Children", *Journal of Political Economy*, 82(2): 111-131.

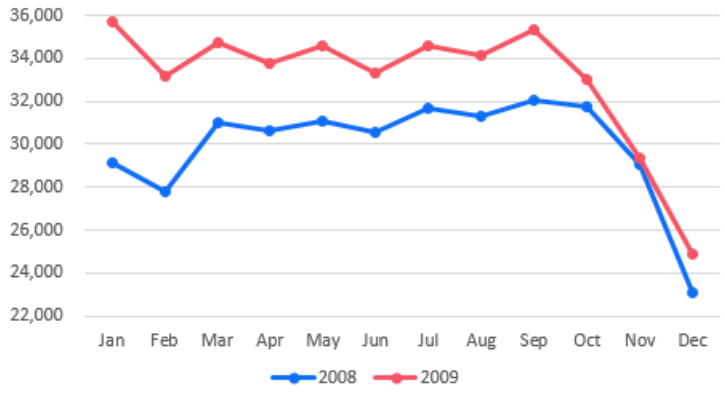
Nollenberger, N. and Rodrguez-Planas, N. (2013). "Full-time universal childcare in a context of low maternal employment: Quasi-experimental evidence from Spain", *Labour Economics*, 36: 124-136.

Voena, A. (2015). "Yours, Mine, and Ours: Do Divorce Laws Affect the Intertemporal Behavior of Married Couples?", *The American Economic Review*, 105(8): 2295-2332.

Appendix A: Evidence of number of births

I complement the birth's evidence coming from the ENAHO with data from the Ministry of Health of Peru (MINSA, for its acronym in Spanish). These data aggregate the number of live births. Ideally, I would have the administrative data in which this report was based, but this is not publicly available and I could not obtain it yet. Of course, without this data I cannot rule out concerns about fine timing of births around the cutoff. However, these aggregate data confirm that there were no spike in births in March compared to April, neither in 2008 nor in 2009.

Figure A1: Number of monthly births (national level)



Source: INEI - MINSA. Statistic report on live births

Appendix B: Variable description

Table A.1 lists and describes the full set of control variables used in the regressions of Sections 4 and 5. Children's controls includes the gender of the child. Mothers' controls include the age of the mother, a dummy indicating whether

the mother is indigenous, a dummy for married or cohabiting mother and dummies for different education levels. Controls for household's structure includes a set of variables indicating the size of the household, the number of members in different age groups and whether there is a male adult living in the household. Regional variables include a dummy variable for each of the four main regions of the country: the coastal region, the hills, the jungle and the Metropolitan area of Lima.

Table A1: Full set of controls

Variable	Description
Child is a boy	Equals 1 if boy, 0 otherwise
Mother's Age	Age of the mother, in years
Indigenous	Equals 1 if first language is native, 0 otherwise
No High School Degree	Equals 1 if has less than High School degree, 0 otherwise
High School Degree	Equals 1 if has High School degree, 0 otherwise
Tertiary	Equals 1 if has Tertiary education degree, 0 otherwise
Some College	Equals 1 if has some College but no degree, 0 otherwise
College and +	Equals 1 if has College degree or higher, 0 otherwise
Married	Equals 1 if married or cohabiting, 0 otherwise
Coastal region	Equals 1 if HH is located in the coast, 0 otherwise
Hills region	Equals 1 if HH is located in the hills, 0 otherwise
Jungle region	Equals 1 if HH is located in the jungle, 0 otherwise
Metropolitan Lima	Equals 1 if HH is located in Metropolitan Lima, 0 otherwise
HH size	Number of households members, total
Number of HH members, 18+	Number of members in the HH aged 18 or more
Number of HH members, 65+	Number of members in the HH aged 65 or more
Number of HH members, 6-11	Number of members in the HH aged 6 to 11
Number of HH members, 12-17	Number of members in the HH aged 6 to 17
Number of HH members, 0-5	Number of members in the HH aged 0 to 5
Number of family members, 0-5	Number of members in the nuclear family aged 0 to 5
Male in the HH	Presence of a male in the HH aged 18 or more
Male in the family	Presence of a male in the nuclear family aged 18 or more