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# ABSTRACT

## The Effects of Paternity Leave on Fertility and Labor Market Outcomes*

This paper studies the effects of a father quota in the parental leave period on households' labor market and fertility decisions. Identification is based on the 2007 reform of the Spanish family benefit system, which extended the sixteen weeks of paid parental leave by two additional weeks exclusively reserved for fathers and non-transferable to mothers. Using a regression discontinuity design, we show that the reform substantially increased the take-up rate of fathers (by as much as 400%), as well as the re-employment probability of mothers shortly after childbirth (by about 11%). However, it did not affect parents' longer-term leave-taking or employment behavior. We also find that the introduction of the two weeks of paternity leave delayed higher-order births and reduced subsequent fertility among older women (by about 15%). These results suggest a limited scope for the father quota to alter household behaviors beyond the parental leave period and reduce gender inequality at the workplace.

**JEL Classification:** J48, J13, J16

**Keywords:** natural experiment, paternity leave, fertility, labor market and gender

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## 1. Introduction

Parental leave policies were initially designed to protect mothers and children's health at an early age, as well as to help balance family and market work. The main economic feature of these policies was the right to return to a previous employment position within a certain period (job-protected leave).[1] However, the labor market effects of family policies are complex and not always well-understood (Olivetti and Petrongolo 2017, Rossin-Slater 2017).

Extended maternity leave permits reduce effective employment immediately after birth, but have negligible or even positive effects on the labor market attachment of mothers in the long-term (Lalive and Zweimüller 2009, Lalive et al. 2014, Schönberg and Ludsteck 2014). However, very generous programs may have a detrimental effect on the wages and job opportunities of women, in particular at the top end of the skill distribution (Blau and Kahn 2013, Albrecht et al. 2015). There is also evidence that extended periods of job-protected leave and financial incentives to increase the take-up rate of mothers can have sizeable positive effects on fertility (Lalive and Zweimüller 2009, Raute 2015).

While in most countries both parents are eligible to take time off from work after a child is born, the large majority of leave-takers are women (OECD 2015). To increase fathers' involvement in childrearing and reduce gender specialization within the family, many countries have implemented parental leave entitlements for fathers which are not transferable to mothers. In a recent survey, Rossin-Slater (2017) reports that "*out of 167 countries with available data, 79 have a paternity leave entitlement in their legislation. Paternity leave is typically much shorter than maternity leave—ranging from one day (...) to 90 days (...)".* By providing a more gender balanced contribution to childcare, time reserved for fathers has the potential to reduce statistical discrimination against women in the labor market if both parents take time off to look after their children.

The existing evidence indicates that a leave period or quota reserved for fathers provides strong short-term incentives to increase their participation in parental leave (Ekberg et al. 2013, Bartel et al. 2015, Patnaik 2015). However, its effectiveness in raising fathers' long-term involvement in childrearing activities and household work is unclear. Ekberg et al. (2013) argue that the lack of fathers' commitment beyond the paternity leave period may respond to the persistence of social norms that public policy

[1] The period of job-absence may be paid, unpaid or partially paid depending on each country specific policy (OECD 2015).

cannot easily modify.[2] A recent survey by Olivetti and Petrongolo (2017) points out that "*While most high-income countries currently have in place leave provisions for fathers, their relatively recent introduction, as well as their more limited take-up rate, imply that the evaluation of their effects (…) is still in its infancy.*"

In this paper, we study the effects of the introduction of a paternity leave quota in Spain. As most Southern European countries, Spain has a relatively low rate of female employment and a not very generous family benefit system. Prior to the reform, the paid parental leave period lasted sixteen weeks, six of them reserved for mothers immediately after birth. While parents had the possibility to share ten weeks of the leave, fathers' participation in parental leave was negligible (see Figure 1). In March 2007, a permit of two weeks exclusively reserved for fathers and non-transferable to mothers was introduced.[3] Figure 1 shows that the new regulation dramatically increased fathers' take-up rate. One year after implementation, 279,756 fathers had taken a paternity leave permit, 54% of all new fathers. In 2016, the share of fathers on paternity leave was around 60% of all births.[4] In January 2017 the parental leave reserved to fathers was extended to four weeks.

We employ a regression discontinuity design to quantify the effects of the introduction of the father quota in March 2007 on parental involvement in childrearing, as well as on the labor market and fertility decisions of households. While the effects on take-up rates and labor market outcomes have been studied for other countries, to the best of our knowledge ours is the first analysis of the effect of paternity leave on fertility.[5]

The empirical analysis suggests that the two weeks period of "use-it or lose-it" parental leave for fathers substantially increased their take-up rate (by 400% of the pre-reform mean), as well as the employment probability of mothers shortly after childbirth (by 11% of the pre-reform mean). However, these short-term responses seem not to

[2] See also Tanaka and Waldfogel (2007), Nepomnyaschy and Waldfogel (2007), Rege and Solli (2013), Huerta et al. (2013), Kluve and Tamm (2013), Dahl et al. (2014), Almqvist and Duvander (2014), Cools et al. (2015), Bünning (2015), Bünning and Pollmann-Schult (2016), Abrahamsen (2017) and Dunatchik and Özcan (2017).

[3] In Spain, the two weeks of father quota represents an 11% of the parental leave period. In Norway, the duration of the father quota is ten weeks which also accounts for the 10% of the total parental leave entitlement, while in Sweden the two-months father quota represents the 20%.

[4] Note that only employed parents can benefit from a paid period of parental leave.

[5] Dahl et al. (2014) examine the effect of male coworkers and brothers take-up rates on fertility and find no measurable effect for Norway.

have had implications for the longer-term employment and leave-taking behavior of parents. These results are consistent with the evidence reported for other countries with more generous family policies (see Ekberg et al 2013 for Sweden or Abrahamsen 2017 for Norway).

In terms of fertility, we document that the policy delayed childbearing and reduced higher-order births among older women (by 15% of the pre-reform mean). This result contrasts with the positive effect on fertility associated with very generous parental or maternity leave policies (see Lalive and Zweimüller 2009 and Raute 2016). We argue that the earlier return to work of mothers eligible for the two weeks quota, combined with a potential increase in childbearing costs for fathers, may have led to extending the time between births, thus negatively affecting the completed fertility of women nearing the end of their fertile cycle.

The short-lasting effect of the reform on parental behaviors and the drop in higher-order fertility cast doubt on the effectiveness of the two weeks gender quota to alter the specialization model or social norms governing family arrangements. It remains to be seen whether the extension to four weeks introduced in January 2017 will be more successful in changing the allocation of parental time to family and labor market activities beyond the leave period.

The paper is structured as follows. In the next section we describe the institutional setting and develop testable hypotheses regarding the effects of the 2007 reform on households' labor market and fertility decisions. Section 3 describes the methodological approach, followed by the description of the data in Section 4. The results are presented in Section 5, and some concluding remarks follow in section 6.

## 2. Background and hypotheses

### 2.1 Institutional setting

As in most countries, parental leave mandates in Spain have evolved to accommodate the increasing number of women in the labor market, with the goal of reducing gender disparities at the workplace. Figure 2 shows the evolution of the employment rate of mothers with children under 1 year old. This rate has more than doubled since the late 1980s (from 25% in 1987 to 60% in 2015). The employment rate of fathers has remained fairly constant, around 90%, over the last few decades, slightly decreasing during the last few years as a result of the economic recession.

As compared to other European countries, the system of family benefits in Spain is rather scarce, though it has become more generous and gender neutral over time. Until 1989, working women were entitled to 14 weeks of paid-protected job leave, and fathers could take 2 days off after the baby's birth.[6] In 1989 the maternity leave period was extended to 16 weeks, with the last 4 weeks transferable to the father.[7]

In 1999, a law to promote work and family life (Law 39/1999) allowed mothers to transfer up to 10 weeks of the parental leave period (16 weeks) to fathers. These weeks could be enjoyed simultaneously or subsequently to those of the mother. The new regulation also established that the first 6 weeks after the baby's birth were compulsory and exclusively reserved for the mother.[8] The introduction and subsequent extension of the shared period of parental leave had negligible effects on fathers' take-up rate.[9]

In March 2007, the Socialist Party in government passed the Law 3/2007 on “effective equality between men and women”.[10] The Law contained measures to foster gender equality in several fields: from electoral norms to violence against women or public contracts. It also modified the parental leave period by introducing a two weeks quota of paid leave exclusively reserved for fathers and non-transferable to mothers.[11] Fathers could take the new leave simultaneously or subsequently to that of the mother, with the only restriction that the 6 weeks after birth were compulsory and exclusively reserved to the mother. The paternity leave period should have been progressively extended to 4 weeks by 2011. However, the severe economic recession in Spain postponed this extension, which only materialized in January 2017. Table 1 summarizes the main contents of the successive parental leave reforms in Spain.

The introduction of the gender quota in 2007 immediately increased the participation of fathers in parental leave (see Figure 1). In what follows we investigate the implications of this reform for the labor market outcomes of parents and their subsequent fertility decisions.

[6] Until 1994 (Law 42/1994, December 30, 1994) female workers on maternity leave were considered as transitory disabled to work and received a 100% wage replacement rate.

[7] The extension of the maternity leave period to 16 weeks was regulated by Law 3/1989. The law was passed on March 3, 1989; published on March 8, 1989 and implemented on March 28, 1989.

[8] The Law 39/1999 was passed on November 5, 1999; published on November 6, 1999 and implemented on November 7, 1999.

[9] In 2006 less than a 2% of the fathers applied for the shared period of parental leave.

[10] The Law 3/2007 was passed on March 22, 2007; published on March 23, 2007 and implemented on March 24, 2007.

[11] Since 1980 fathers were entitled to 2 days of job absence after the baby's birth. The 2007 reform introduced a 13 days period of leave with a 100% wage replacement rate.

## 2.2 Expected effects on labor market and fertility decisions

The main goal of reserving a share of the parental leave for fathers is to balance the distribution of childrearing costs within the family, achieve a less gender specialized home production model and foster the labor market prospects of women.

In the short-run, the introduction of a two weeks "use-it or lose-it" permit for fathers is expected to increase their participation in parental leave. These two additional non-transferable weeks may also affect the return-to-work decision of mothers, as they can go back to work earlier while the father stays home with the baby, potentially delaying the start of paid childcare. These two effects should be particularly strong if families exhausted the shared period of the parental leave before the reform (Patnaik 2015).

If the gender quota affects father's participation in childrearing activities beyond the parental leave period, there may be related longer-run effects. A more egalitarian distribution of childcare chores may allow women to devote more time to paid work and pursue less gendered professional careers (Becker 1991, Lazear and Rosen 1990). In addition, less specialization in home production may reduce statistical discrimination. Employers may anticipate that fathers will take time off to care for their children and modify their hiring and promotion decisions in favor of female workers (Phelps 1972, Lazear and Rosen 1990). Therefore, a family policy that shifts part of the childcare burden to fathers may positively affect women's longer-term labor market outcomes in terms of employment participation, career achievements and life time earnings. Conversely, it may worsen the labor market prospects of fathers if their participation in home production increases. Indeed, a male wage penalty related to parental leave absences have been documented for some countries (Albrecht et al. 1999, Bertrand et al 2010).

The introduction of a reserved share of the parental leave for fathers may also affect households' fertility decisions. The policy reduces the direct cost of having children by offering two additional weeks of paid parental leave. It may also represent an additional decrease in the direct cost for mothers if fathers become more involved in childrearing. Conceivably, after the reform women may be more willing to have (more) children, while the opposite would be true for men, assuming both prefer lower involvement in childcare. The expected effect of a shift in childcare costs is ambiguous and depends on the distribution of men’s and women's preferences over fertility and childcare activities (Doepke and Kindermann 2016).

At the same time, the opportunity cost of childrearing for women may increase if the quota reduces the duration of career interruptions and foster their labor market prospects (Adda et al 2017). As a result, fertility will decrease provided that fathers' opportunity cost does not decrease to the same extent.

In what follows, we estimate the effect of the gender quota on the leave-taking, employment and fertility decisions of households, and investigate the potential mechanisms driving the results.

## 3. Methodology- A natural experiment

We follow a regression discontinuity approach and focus on families (mothers and fathers) who had a child right before and right after the implementation of the reform (between 3 and 9 months before and after March 24, 2007). The "running variable" is the month of birth,[12] and the treatment variable is a dummy indicating the "post-reform" period (families who had a child after March 2007). Note that none of the other reforms that were implemented simultaneously with the extended paternity leave (i.e. electoral norms, regulation of gender violence, etc...) would be expected to affect differentially families with children born right before and right after March 2007.

We estimate the following equation:

$$(1)\ Y_{it} = \alpha + \beta Post_{it} + \gamma f(m_{it}) + \lambda Post_{it} * f(m_{it}) + \mu X_{it} + u_{it},$$

where $Y$ is an outcome variable (say, mother's labor market participation 6 months after birth) for household $i$ who had a child in month $t$, *Post* is an indicator for the child being born after the introduction of the gender quota, $m$ is the running variable (taking value 0 for couples who had a child in April 2007, the first month after the reform was implemented, -1 for the month before, 1 for the month after, and so on), and $X$ is a set of demographic controls, such as age, education and immigrant status. The $f(.)$ function is a polynomial on $m$.

The two assumptions driving identification in equation (1) are: i) All other relevant factors at the time of the treatment vary smoothly at the cutoff; and ii) There is no selective sorting across the RD threshold. We provide evidence in support of these two assumptions in the results section ("Validity checks"). We show balance in

[12] When available, we use the exact date of birth instead.

covariates in support of i), and evidence that there was no bunching in the number of births after the cutoff in support of ii).

In terms of outcomes, first we examine the leave-taking behavior of fathers and mothers right after birth and their short-term labor market responses. Then we look at longer-term effects (up to four years after the implementation of the reform) on labor market, parental involvement and fertility outcomes.

## 4. Data and descriptive statistics

In the empirical analysis we use three different data sets. Most of the results are obtained from the Spanish Labor Force Survey (*Encuesta de Población Activa*), which contains information on labor market (and fertility) outcomes of households. We supplement the analysis with administrative Social Security data (*Muestra Continua de Vida Laborales*), with information on earnings and employment status. Finally, we also use administrative birth-certificate records to validate our fertility results.

The Spanish Labor Force Survey is conducted every quarter, with the main purpose of collecting data on the labor market status of individuals and households. About 65,000 households are interviewed each quarter, including approximately 180,000 people. We define a sample of families with children born before (control group) and after (treatment group) the date of implementation of the reform (March 24, 2007). Since the Labor Force Survey does not contain information on the exact date of birth, but only month and year, the treatment group includes families with children born from April 2007 onwards. In our main sample we compare families with children born in a span of 12 months around the reform date (i.e. from October 2006 to March 2007 in the control group, and from April 2007 to September 2007 in the treatment group).[13] We restrict the sample to families where the father and the mother both live in the household.[14]

Since the Labor Force Survey is not longitudinal, we do not observe all families at the same point in time after childbirth. Therefore, to conduct the empirical analysis we define a time window for the interviews (e.g., all quarters in 2008) and analyze the behavior of families with children born close to the introduction of the reform, who

[13] We investigate the robustness of our results to reducing and expanding the span to 6 and 18 months, respectively.
[14] Notice that the policy under study aims at increasing father's involvement in childcare activities as an instrument to reduce the impact of childbearing on women's career. Thus the policy change is expected to have mainly affected heterosexual couples.

were interviewed during the specified quarters. As a result, children in the treatment group are, on average, observed at a younger age than children in the control group. While in the empirical analysis we account for this compositional effect by including a polynomial in month of birth, the comparison of the descriptive statistics between the pre and post-reform groups reflects a mixture of the effects of the reform and the child-age effect.

To measure the short-term effects of the reform, we observe the relevant families shortly after the birth of the child. We define a first sample of families who were interviewed by the Labor Force Survey in the last quarter of 2006 or in any quarter in 2007. Children born right around the policy change, in March-April 2007, are on average 4 months old in this sample, which we will thus refer to as *"child age 4 months"*. A second sample is defined by adding interviews from the first two quarters in 2008, so that the average age of children born at the threshold in this sample is 8 months. We will refer to this sample as *"child age 8 months"*.

The longer-term effects of the policy are measured on families observed in the last two quarters in 2008 or in any quarter in 2009, when children born at the threshold are, on average, 24 months old. This is the *"child age 24 months"* sample. The polynomial in month of birth in equation (1) controls for age in all specifications.

We estimate the effects of the introduction of the gender quota on the behavior of both mothers and fathers in terms of participation in parental leave, involvement in childrearing, labor market outcomes, and fertility decisions. Table 2 presents the descriptive statistics for the main labor market and fertility outcomes.

The Labor Force Survey questions individuals about their labor market status in the week prior to the interview. The survey does not ask about parental leave specifically, but it does indentify individuals who are temporarily absent from work due to parental leave (or other reasons) during the week of reference. Since parental leave in Spain is relatively short, particularly for fathers, and we observe leave-taking in only a single week, we will miss most of the leave episodes. However, as long as births and the average length of leave are uniformly distributed throughout the year, the percentage change in leave-taking estimated to result from the policy will be accurately captured, although the level will be understated (Bartel et al. 2015).

Table 2 compares the outcomes of parents of children born within 6 months before and after the reform in the short-term (samples *"child age 4 months"* and *"child age 8 months"*) and the longer-term (sample *"child age 24 months"*). The table suggests that the policy may have been effective in increasing the short-term incentives for fathers to take parental leave. In the sample of families observed shortly after birth (*"child age 4 months"*) the average take-up rate before the reform was 0.3%, which increased to 3.1% immediately after implementation. In the *"child age 8 months"* sample, where families are observed a few additional months after childbirth, the rates are 0.2% before and 1.6% after the policy change. The comparison in terms of the other labor market outcomes does not suggest important differences across the two groups of fathers.

Table 2 also indicates that the take-up rate of maternity leave is higher in the post-reform group (29% in *"child age 4 months"* and 15% in *"child age 8 months"*) than in the pre-reform group (14% and 9%, respectively). The introduction of the two weeks gendered quota is not expected to have a first-order effect on the leave-taking behavior of mothers. The differences in take-up rates of mothers may be driven by the age composition of children in the control and the treatment group and the fact that mothers in the treatment group in the samples *"child age 4 months"* and *"child age 8 months"* are more likely to be observed right after child birth, during the 16 weeks of parental leave. The empirical analysis below allows us to assess whether these differences persist once the child-age effect is taken into account.

For mothers, the table also reveals some important differences in terms of other labor market outcomes. In the post-reform group, mothers are more likely to work shortly after birth (59% in *"child age 4 months"* and 56% in *"child age 8 months"*) than those in the pre-reform group (52% and 53%). However, these differences do not persist in the longer term (*"child age 24 months"*).

The second panel in Table 2 also suggests that the policy may have had effects on fertility. The comparison of the higher-order birth rates between pre and post-reform mothers suggests a decrease in fertility after the introduction of the gender quota. The probability of additional births decreases from 13% for pre-reform mothers to 7.7% for post-reform mothers.

We supplement the main analysis with information on earnings obtained from the *Continuous Sample of Working Lives* (*Muestra Continua de Vida Laborales*), a large micro-level panel data set assembled by the Spanish Social Security Administration. The information is obtained from a 4% random sample of all individuals that are (or have been at some point in the reference year) affiliated with Social Security. We use the 2009 sample, and select the subsample of adult men and women living in a household with a child born close to March-April 2007. We then add up all earnings during the 6 and 12 months following the birth of the relevant child (third panel of Table 2). Fathers' earnings in the pre and post-reform groups are similar both 6 months and 12 months after having the child (around 8,700 and 17,590 Euros respectively). However mothers' earnings are slightly higher in the post-reform group both 6 months (5,681 versus 6,191 Euros) and 12 months after birth (11,951 versus 12,681 Euros). For mothers, the employment probability 6 months after birth is also higher for the post-reform group (60% versus 64%). Finally, the data also includes information on unpaid parental leave.[15] The descriptive statistics show an important decrease in the probability of being on unpaid leave among post-reform mothers ( 7% versus 5%).

Finally, the fertility analysis is supplemented with information from birth-certificate data. These micro data are publicly available through the National Statistical Institute, and they record information for the universe of children born (and registered) in Spain annually. We use information on the date of birth of each child, as well as the date of birth of the previous child, to construct subsequent birth rates for women having a child around March 2007. We construct the fraction of women who had a child in the following 2 and 4 years, by exact date of birth of the child born close to the introduction of the father quota. Since we observe the exact date of birth, we can focus on a narrow window of birth-dates around the threshold (between 1 and 11 weeks). On average, about 6-7% of women had another child during the two years after the 2007 one, while almost 24% had at least one additional child within four years.

## 5. Results

The identification strategy relies on comparing the behavior of families who had a child right before versus right after the introduction of the two weeks father quota on March 24, 2007. Thus, for all labor market and fertility outcomes identification is achieved via a Regression Discontinuity Design (RDD), as formalized in equation (1). We supplement the analysis with Difference-in-Difference specifications (DiD) to account for potential month-of-birth effects. We first provide some validity checks in support of our identification assumptions, discussed in section 3 ("Methodology").

[15] In Spain parents are entitled to unpaid parental leave until the child is three years old. The job is protected for the first year of leave only. After the first year, job protection is restricted to a job of the same category. The Social Security data do not contain information on paid parental leave.

## 5.1 Validity checks

One key identifying assumption in the RDD is that parents do not time births or change eligibility status in response to the policy. If this assumption holds, the distribution of pre-determined characteristics of parents should not abruptly change around the reform date. In order to test it, we run the RD regression in equation (1) with the parental characteristics (i.e. age, education, immigrant status, etc.) as dependent variables. The estimates are reported in Table 3. Most of the covariates do not exhibit significant jumps at the cutoff point. We should highlight however the higher presence of immigrant mothers among the pre-reform group. We will thus control for immigrant status in all of our specifications.

We conduct a second test for the possibility that (some) families were able to time birth after the threshold in order to take advantage of the extra weeks of paternity leave. We estimate equation (1) aggregated at the date of birth level, where the outcome variable is the daily number of births (as reflected in the birth-certificate data). The running variable is the exact date of birth. The coefficient of interest would capture any discontinuity in the daily number of births right at the threshold. The results are presented in Table 4. We find no evidence of bunching at the threshold, suggesting that families did not selectively sort across the cutoff date for eligibility.

## 5.2 Effects on fathers

The reform approved in March 2007 extended the 16 weeks of the parental leave period by 2 weeks, exclusively reserved to fathers and non-transferable to mothers. The aim of the policy was to introduce an incentive for fathers to participate in parental leave. Thus, we first investigate the effect on their take-up rate. Figure 3 shows the fraction of fathers by month of birth of the child who reported being on parental/paternity leave at the time of the survey interview, in the sample *"child age 8 months"* (i.e. interviews conducted between the third quarter in 2006 and the second quarter in 2008) of the Labor Force Survey. The discrete jump right after March 2007 is clear.

Table 5 shows the corresponding regression results. The table displays the coefficient on the post-reform indicator ($\beta$) in equation (1). The estimates in column (1) are obtained from the comparison of fathers with children born within a 6-month window before and after the reform.

The estimated coefficients confirm that the policy substantially increased the take-up rate of fathers. In the short-term, the introduction of the two weeks quota led to a 0.9-1.3 percentage point increase in fathers' leave-taking probability. These point estimates are highly statistically and economically significant, representing more than 400% of the pre-reform average participation rate of 0.2-0.3%. These effects persist when we compare the leave-taking behavior of fathers with children born within a 3-month window around the reform in column (2), or within a 9-month window in column (3).

The estimates of the longer-term effect obtained from the sample *"child age 24 months"* are smaller in magnitude, and only significant when the comparison window is restricted to 6 months around the reform. This lack of persistence in leave-taking behavior is consistent with the design of the parental leave system in Spain. Accordingly, fathers have to be on leave simultaneously or immediately after the 16 weeks of maternal leave. Thus no long-term effects on parental leave participation are to be expected.

In columns (4) to (6) we investigate whether the reform affected the probability of fathers to be on leave for family reasons other than parental leave (e.g. care of sick children).[16] The estimates on the post-reform indicator are only statistically significant, at the 10% level, in the short-term sample when the comparison window is 3 months. These results suggest that the policy had a limited effect in increasing fathers' participation in childrearing activities beyond the non-transferable period of parental leave.

The remaining columns in Table 5 ((7) to (9)) display the estimated effect on the employment rate of fathers with children born within 6, 3 and 9-months around the reform. None of the coefficients on the post-reform indicator is statistically significant. We have also explored whether the reform affected the intensive margin of fathers' labor supply, as well as other labor market outcomes. The results, available upon request, do not reveal any effect on hours worked, occupation, or contract type.

The estimates in Table 5 are obtained from comparing the behavior of fathers with children born in different seasons. Children in the control group were born between October 2006 and March 2007, while those in the treatment group were born between April 2007 and September 2007. Previous studies show that there are seasonal changes

[16] Ekberg et al (2013) employ the probability to be on leave to take care of sick children as the measure of involvement in household work.

in the characteristics of women giving birth throughout the year (Clarke et al. 2016, Buckles and Hungerman 2008). In the presence of marital sorting on the basis of those characteristics, fathers of children born in different seasons may also be different. To incorporate these concerns, we also estimate specifications that include multiple birth years and control for "seasonality" by including calendar month of birth fixed effects in equation (1), thus supplementing the regression discontinuity design (RDD) with difference-in-difference (DiD) estimates. In particular, we include in estimation fathers with children born 18 months before and after the reform (between October 2005 and September 2008).

The results of the DiD estimates are displayed in columns (1) and (2) in Table 6. To conduct the difference-in-difference estimation we need to modify the timing of our samples. We construct a first sample (DiD1) where fathers with children born 18 months around the reform are interviewed in the third or fourth quarter in 2005, or in any quarter in 2006, 2007 and 2008. We also employ a second sample that extends the previous one with the observations in 2009 (DiD2), and a third one that adds those in 2010 (DiD3). The results in Table 6 confirm that our findings are not driven by seasonality. The estimated effect on the take-up probability is statistically significant and of similar magnitude to that estimated in Table 5. Note that the size of the effect decreases as the span of the sample increases, since fathers are less likely to be on leave as the child ages. The estimates in column (2) of Table 6 also confirm that the policy did not affect the employment outcomes of fathers.

The Labor Force Survey does not contain information on wages or earnings. Thus we rely on the Social Security data to assess the impact of the quota on earnings. We employ again the specification in equation (1), and replace the dependent variable with a measure of earnings, either during the 6 or 12 months following birth. The estimates in Table 7 suggest no effect on subsequent earnings for fathers. The coefficients are small, switch signs across specifications, and are mostly not statistically different from 0. The analysis with Social Security data also confirms the null effect on employment. We also find that fathers affected by the reform were not more likely to be on unpaid parental leave in the months following birth.[17]

[17] In Spain, both mothers and fathers can be on unpaid job-protected parental leave until the child turns 3 years old.

Opponents to the introduction of father quotas argue that the provision of incentives to participate in parental leave would damage fathers' work careers as they would now invest more time to childcare activities. Previous evidence for Sweden suggests that while the one-month quota had an immediate positive effect on fathers' parental leave participation, it affected neither their employment prospects nor the distribution of household tasks (Ekberg et al. 2013). Thus it should not come as a surprise that the (shorter) quota in Spain did not affect the employment outcomes of fathers or their longer-term involvement in childrearing, despite the important increase in parental leave participation shortly after birth.

### 5.3 Effects on mothers

We next investigate the effect of the reform on the leave-taking behavior and employment of mothers. Table 8 presents the same structure as Table 5. Columns (1), (4) and (7) display the results for our preferred specification, where mothers of children born up to 6 months before the reform are compared to those born up to 6 months after. Neither in the short- nor in the longer-term does there seem to be any effect of the policy on the parental leave-taking behavior of mothers. This result is confirmed in the alternative specifications in columns (2) and (3) when the time window for comparison is 3 and 9 months respectively. The probability to be on leave for other family reasons also appears unaffected (columns (4) to (6)). The absence of effects on the leave-taking behavior of mothers is consistent with the fact that the reform did not affect the duration or generosity of the leave available to them, neither did it seem to increase fathers' involvement in childrearing activities beyond the parental leave period.

In contrast, the two-additional weeks of parental leave do seem to have affected the employment of mothers in the short-run. The estimates in column (7) indicate that post-reform mothers were about 6 percentage points more likely to be employed a few months after giving birth (in samples *"child age 4 months"* and *"child age 8 months"*) than pre-reform mothers, which represents an 11% increase from the pre-reform baseline employment rate of 52.5%. This finding is confirmed in columns (8) and (9) with the time window around the reform fixed at 3 and 9 months.

Increasing the labor market attachment of mothers was an intended effect of the policy. The two weeks of parental leave reserved for fathers should have helped mothers to smooth their transition into employment after childbirth. Fathers are allowed to take these two weeks subsequently or simultaneously with the mother. Some critical views argue that fathers tend to enjoy their period of leave jointly with mothers, reducing the effectiveness of the policy. Table A1 in the Appendix investigates this possibility by estimating the effect of the gender quota on the joint leave-taking decisions of co-habiting spouses. The results show that the reform significantly increased the father-only leave-taking probability by 0.6 to 1.1 percentage points, while the other alternatives (i.e. mother-only or both) remained unaffected. The magnitude of this effect is similar to that obtained when the leave-taking behavior of fathers is separately analyzed in Table 5. These findings reinforce the view that the quota raised the take-up rate of paternity leave and allowed mothers to return to the workplace earlier after birth.

Going back to Table 8, the estimates in the last row *"child age 24 months"* indicate that the positive short-term effect of the quota on mother's employment was not long-lasting. None of the estimated coefficients on the post-reform indicator is statistically significant two years after birth in any of the alternative specifications.[18] The absence of an effect on mothers' employment beyond the first year likely responds to the limited scope of the quota in altering the distribution of childrearing tasks beyond the parental leave period. [19]

We turn to the Social Security records to investigate whether the reform had any effect on mothers' earnings prospects. The results are reported in Table 9, for four alternative specifications and three different samples that vary in the length of the window around the reform. The first column suggests that women who had a child after the introduction of the gender quota earned significantly more during the first year after birth. The magnitude of the coefficient is about 5% of pre-reform average annual earnings (see Table 2). However, once we widen the window around the threshold and control for a linear trend in month of birth, the results become smaller in magnitude and insignificant. The results in the third columns suggest that earnings during the 12 months after birth were about 3% higher for women in households that were eligible for the extra weeks of paternity leave, but the coefficient is not precisely estimated, and is only significantly different from zero with 90% confidence.

The results for employment after 6 months offer some support for the previous result that mothers affected by the policy change had higher employment rates shortly

[18] We have also investigated the effect of the quota on the intensive margin of mothers' labor supply and other labor market outcomes. Only the probability of being self-employed increased by 2 percentage points, which represents a 50% increase with respect to the pre-reform mean of 4%.

[19] The results on mothers' take-up behavior and employment are robust to controlling for seasonality (see Table 6).

after birth. Finally, we find that affected mothers were less likely to take extended unpaid parental leave from the job they held at the time of birth. This probability decreases by about 2 percentage points, which represents a decrease of 27% with respect to the pre-reform mean of 7.3%.

In sum, these results indicate that a policy that reserves time of the parental leave exclusively for fathers allows mothers to return to work earlier after childbirth and reduces the incentives to be on extended unpaid leave. However, this earlier return has a limited effect on their longer-term work careers in terms of earnings and other labor market prospects.

### 5.4 Heterogeneity

Our investigation so far reveals a substantial effect of the quota on the parental leave take-up rate of fathers and the return-to-work decision of mothers. We next investigate the presence of heterogeneity in these responses across workers with different characteristics.

Time off from work may have a higher opportunity cost for high-skilled (more educated) workers. In Spain, the parental leave period is fully paid and the introduction of the quota does not seem to affect longer-term labor market prospects. However, previous studies in other countries have documented a parental leave wage penalty particularly strong among men. Albrecth et al. (1999) argue that this penalty responds to a signal interpreted by the employer as a lack of commitment with the professional career. Accordingly more skilled workers may be less prone to take time out due to their higher costs in terms of employment prospects. Another reason that would lead to a lower take-up rate by more skilled workers is their better ability to substitute parental time at home by formal childcare due to their higher incomes. Therefore, we expect the take-up rate of high-skilled workers to be less responsive to the reform.

Table 10 shows the effects of the policy on the sample of parents with children born within 6 months around the reform, separately by their level of education. The comparison of the estimates between college and non-college graduates indicates that our previous findings are largely driven by less skilled workers. In particular, the take-up rate of college graduate fathers does not respond to the introduction of the two weeks quota, while that of the non-college increases by 1 percentage point, similar to the estimate obtained from the whole sample in Table 5. At the same time, only the employment probability of less skilled mothers increases after the reform. The magnitude of this increase is slightly larger than that obtained for the whole sample in Table 8. This finding is graphically illustrated in Figure 4.

## 5.5 Fertility effects

As discussed in Section 2.2, the expected effect of reserving part of the parental leave for fathers on fertility is ambiguous. If the quota increases father's participation in childrearing, the direct cost of having children for women may decrease, so they may want to have more children (although fathers may want to have fewer). In contrast, a greater involvement of eligible women in labor market activities may increase their opportunity cost of having additional children.

Figure 5 displays the effect of the reform on subsequent fertility. The graph clearly shows a drop in higher-order births after the introduction of the gender quota. Table 11 quantifies this drop using data from the Labor Force Survey on the sample of women who had children within 6 months around the reform. Subsequent fertility is measured by the probability of having an additional child up to three years after the reform in March 2007 (i.e. 2008, 2009 and 2010). The first row in Table 11 indicates a substantial decrease in fertility among affected mothers of about 2 percentage points, which represents 15% of the pre-reform mean probability (13%).

Further inspection of the estimates in Table 11 reveals the presence of heterogeneity in the response by different groups. Consistent with the reported results on leave-taking and employment, the effect is only significant among non-college graduates. For this group, subsequent fertility decreases by about 2.4 percentage points. The effect of the reform also varies across age groups. Fertility among women older than 35 decreases by about 3 percentage points, while that of younger women remains unaffected.

This age asymmetry suggests that delayed childbearing among eligible women may be responsible for the drop in higher-order births. The medical literature shows that the decision to postpone childbearing is more costly for the completed fertility of women nearing the end of their fertile cycle (Schmidt et al. 2012). The birth-certificate data allows us to further investigate the impact of the reform on postponed fertility. Figure 6 graphically shows the effect of the reform on birth spacing. The sample includes women having a child within 12 weeks around the introduction of the reform (March 24, 2007), who had another child by the end of 2013 (more than 6 years later). The vertical axis in Figure 6 displays the average number of days between the reference birth and the following one. The graph shows that families who were eligible for the two weeks of paternity leave took longer to have their next child.

To quantify this effect, we estimate equation (1) on the sample of women giving birth within a few days (or weeks) of the reform date. The dependent variable is the number of days between the birth that took place close to March 2007 and the next birth (within the following six years). Since the birth-certificate data include the exact date of birth (i.e. day, month and year), we replace the linear trend in month ($m$) by a linear trend in date of birth, and interact it with the post reform indicator. The main results are reported in Table 12. Each column includes a different number of birth-dates (between 1 and 10 weeks) around March 24, 2007.[20] The estimates indicate a statistically significant increase in birth spacing in most specifications. On average, families eligible for the father quota postponed the following birth by 17 to 27 days.

Finally we test the robustness of our fertility results using the birth-certificate records. In Table 13 each column also includes a different number of birth-dates (between 1 and 10 weeks) around the threshold. The table shows the estimates for the probability of having an additional child in the following 2 years after the introduction of the reform.[21] The upper part of the table reports the results for older women, while the lower part shows the results for women younger than 32. The estimates confirm the findings in the Labor Force Survey. Older women eligible for the two weeks of the paternity leave are between 10 and 15% less likely to have an additional child during the following two years, while the fertility of younger women remain unaffected. After four years, the magnitude of the estimated coefficient is similar, although the coefficients are estimated less precisely (see Table A2).

On the whole, our estimates indicate a drop in subsequent fertility among older women in response to the introduction of the father quota. It seems likely that the earlier return to work of eligible women and a possible increase in childrearing costs for fathers resulting from their participation in paternity leave may have led couples to extend birth spacing. This may have negatively affected the subsequent fertility of older women, with a more limited remaining time to conceive.

[20] Note that Table 12 compares the future fertility decisions of couples with the relevant birth very close to the threshold (1 to 10 weeks). This short time comparison window prevents changes in the composition of families in response to the policy.

[21] The table A2 in the Appendix shows the estimates for the probability of having an additional child in the following 4 years after the introduction of the reform.

The results of our study confirm previous findings on the limited effects of the father quota on the leave-taking and labor market outcomes of eligible households. However, the negative effect on fertility was not documented in previous studies and goes in the opposite direction of the effects for parental or maternity leave extensions. The delay in higher-order births resulting from the introduction of the father quota may have undesirable effects for the completed fertility of countries where parenthood starts relatively late.

## 6. Conclusions

We investigate the effects of the introduction of two extra weeks of paid parental leave exclusively reserved for fathers and non-transferable to mothers, from a baseline where fathers were only entitled to 2 days of paid leave after the birth of a child. Using evidence for Spain, we find that the two week quota substantially increased the short-term incentives for fathers to participate in parental leave. We also find a significant increase in the re-employment probability of mothers shortly after childbirth. Finally, we uncover a negative effect of the reform on higher-order fertility, most likely driven by the decision to postpone childrearing among older couples.

The short-term nature of our results on the employment and leave-taking behavior of parents is consistent with the evidence in Ekberg et al. (2013), where the one month father quota in Sweden was found to have negligible long-term effects on fathers' involvement in childrearing activities. These results suggest that there may be a limited scope for public intervention to alter the social or cultural norms that govern the persistent gendered division of household tasks. It could also be that the size of the quota is not long enough as to motivate a change in individual behavior in the household and at the workplace. A more intensive reform, such as equalizing the duration of parental leave across both genders may have a stronger effect on the division of childcare chores.

We also find evidence that the introduction of a gender quota in the parental leave may have undesired effects on fertility. This may be particularly worrisome in countries with lower-than-replacement fertility rates such as Spain. The low incidence of the father quota on the promotion of gender equality and its potential negative effect on fertility suggest that other policies, such as subsidized childcare, may be more effective in balancing the allocation of paid work across genders (Olivetti and

Petrongolo 2017). For example, flexible and publicly available childcare would allow men and women to choose the optimal amount of time devoted to labor market activities by outsourcing home production. More research is needed to identify the optimal combination of family policies to achieve gender equality in the labor market and at home.

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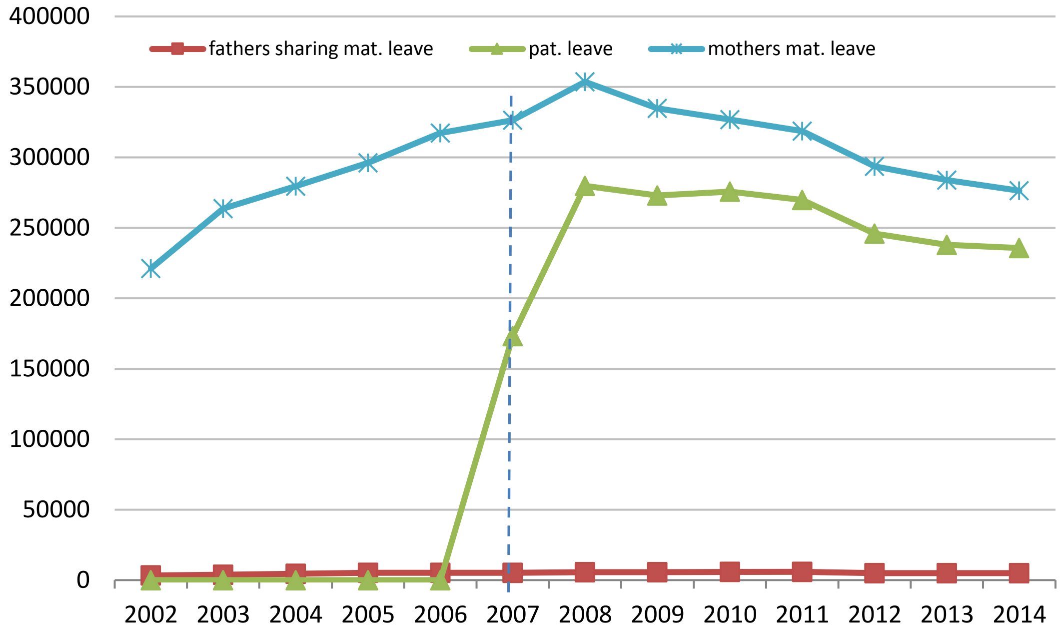
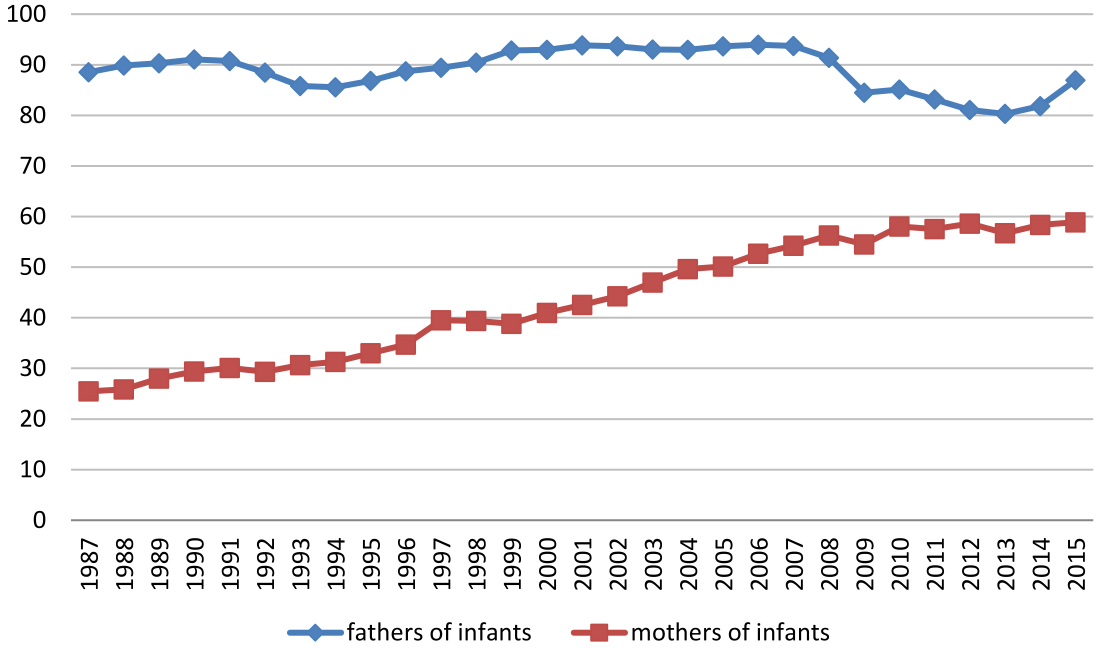
## Figures

Figure 1: Number of mothers and fathers taking maternity/paternity leave



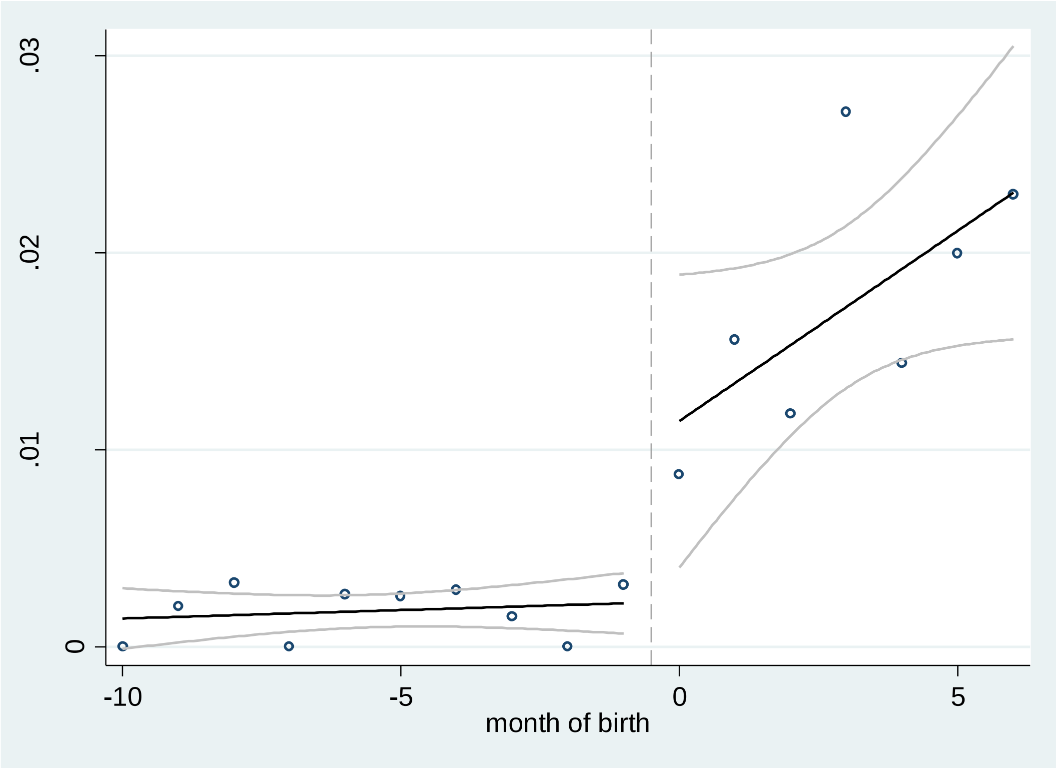

Source: Administrative data from Social Security records.

Figure 2: Employment rates of mothers and fathers of infants (less than 1 year old)



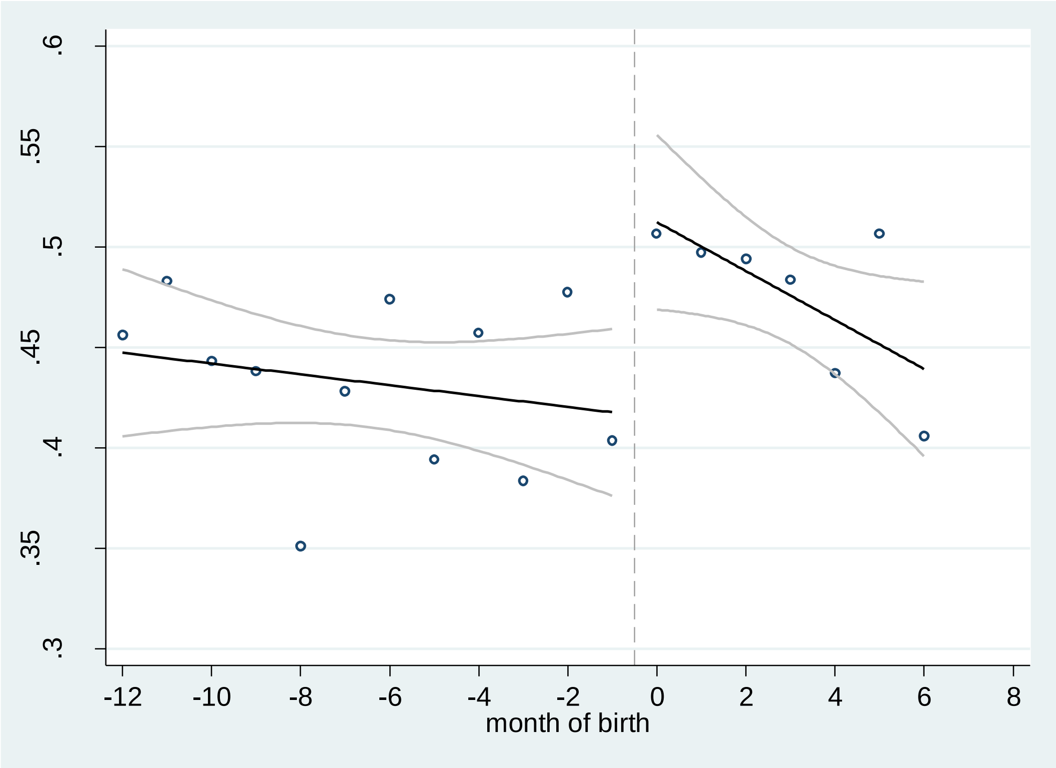

Source: Spanish Labor Force Survey (EPA). Annual employment rate for the period 1987-2015.

Figure 3: Fraction of fathers on parental leave, by month of birth



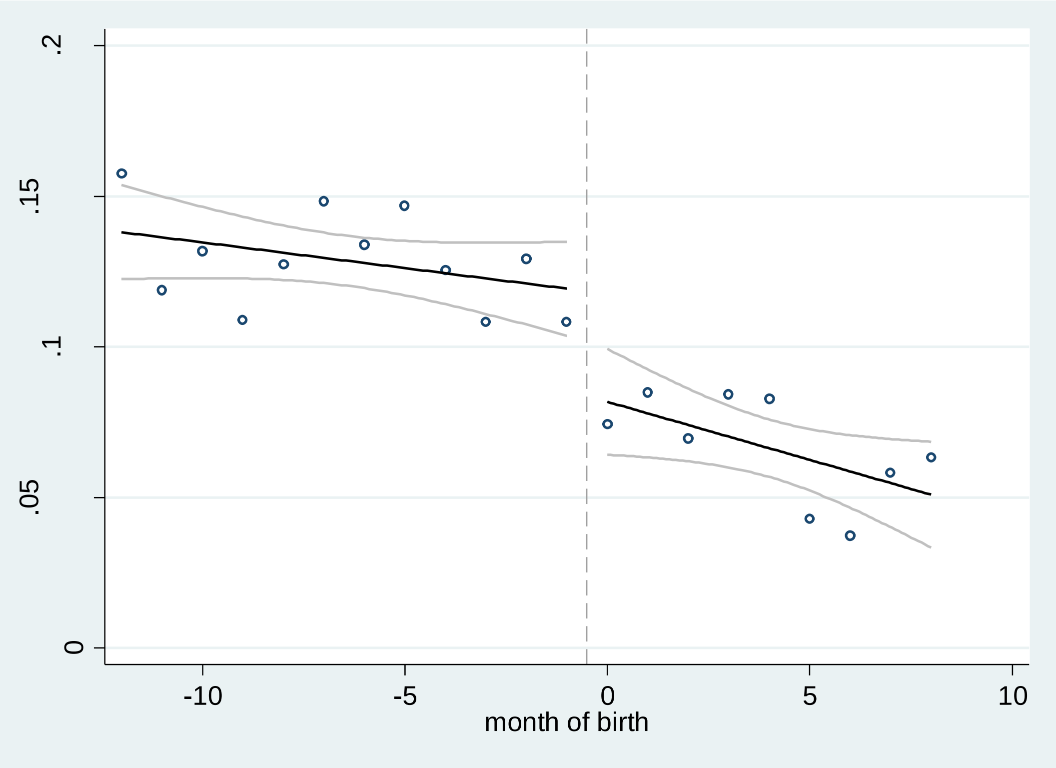

Note: Labor Force Survey data. The horizontal axis is the month of birth of the child, where 0 is April 2007, 1 is May 2007, and so on. The vertical axis is the fraction of fathers aged 16 to 50 living with a partner who report to be on leave from their job at the time of the survey in the sample *"child age 8 months"* (i.e. interviewed between the third quarter in 2006 and the second quarter in 2008).

Figure 4: Employment rate of mothers (about 4 months after birth), by month of birth



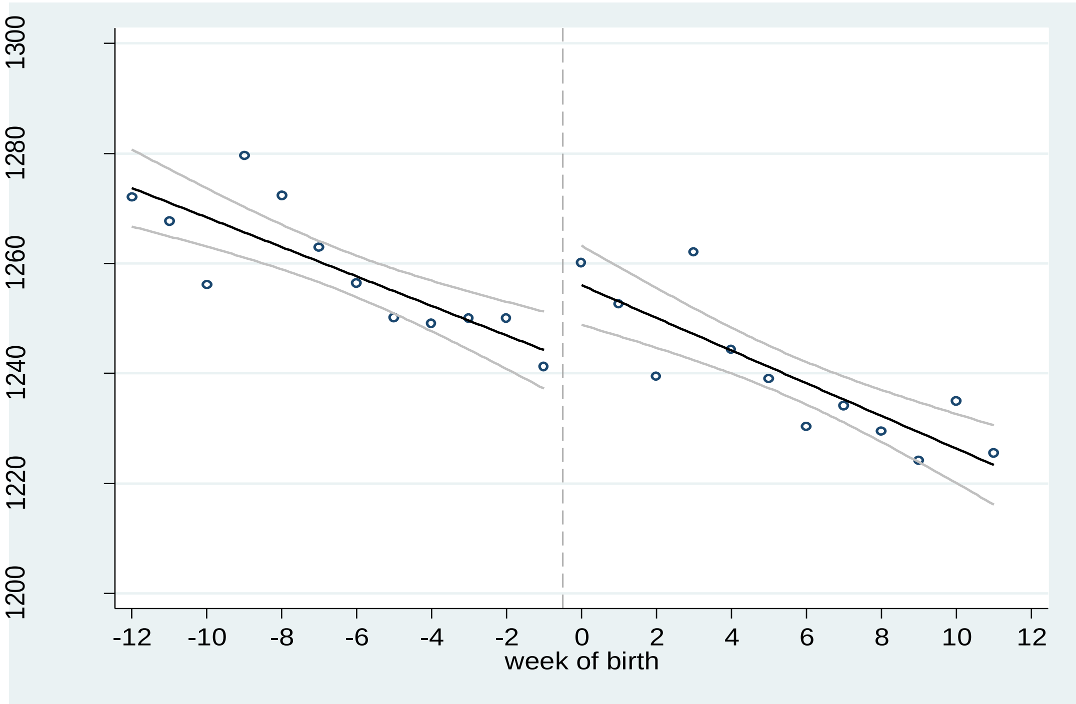

Note: Labor Force Survey data. The horizontal axis is the month of birth of the child, where 0 is April 2007, 1 is May 2007, and so on. The vertical axis is the employment rate of mothers aged 16 to 50, with no university education, and living with a partner at the time of the survey, in the sample *"child age 4 months"* (i.e. interviewed between the third quarter of 2006 and the last quarter of 2007).

Figure 5: Subsequent fertility rate by month of birth




Note: Labor Force Survey data. The horizontal axis is the month of birth of the child, where 0 is April 2007, 1 is May 2007, and so on. The vertical axis is the faction of women aged 16 to 50 living with a partner at the time of the survey who had another child in 2008, 2009 or 2010.

Figure 6: Birth spacing by week of birth




Note: Birth-certificate records. The horizontal axis is the week of birth of the child, where 0 is the week starting in March 24, 2007. The vertical axis is the average number of days between the birth close to the threshold and the following birth to the same mother (for the sample of mothers having another child by the end of 2013).

## Tables

Table 1: Parental leave reforms in Spain

| | ***March 1980***<br>Statute of Rights for Workers<br>***March 1984***<br>Law 30/1984 for the reform of the Public Service | ***March 1989***<br>Law 3/1989 to extend maternity leave to 16 weeks and to promote gender equality at the work place | ***November 1999***<br>Law 39/1999 to promote work and family life | ***March 2007***<br>Law 3/2007 on effective equality between men and women |
|---|---|---|---|---|
| **Fathers** | 2 days of paid job absence after the baby's birth | 2 days of paid job absence after the baby's birth | 2 days of paid job absence after the baby's birth | 2 days of paid job absence after the baby's birth<br><br>13 days of job protected paid leave (non-transferable to the mother) (1) |
| **Mothers** | 14 weeks of job protected paid leave (non-transferable to the father) | 16 weeks of job protected paid leave. The first 6 weeks after birth are compulsory and exclusively reserved to the mother, the last 4 weeks can be transferred to the father | 16 weeks of job protected paid leave. The first 6 weeks after birth are compulsory and exclusively reserved to the mother, the other 10 weeks of the leave can be transferred to the father and enjoyed simultaneously or subsequently to that of the mother | No change |

(1) The paternity leave period was extended to 4 weeks in January 2017.

Table 2: Descriptive statistics

A. Labor market outcomes - Labor Force Survey (EPA)

| | N.obs. Pre-reform | N.obs. Post-reform | On parental leave Pre-reform | On parental leave Post-reform | Employed Pre-reform | Employed Post-reform |
|---|---|---|---|---|---|---|
| **Mothers** | | | | | | |
| Sample : *Child age 4 months* | 2,562 | 1,373 | 0,143 | 0,286 | 0,525 | 0,590 |
| Sample: *Child age 8 months* | 3,970 | 2,764 | 0,094 | 0,152 | 0,532 | 0,561 |
| Sample: *Child age 24 months* | 4,215 | 4,313 | 0,011 | 0,011 | 0,577 | 0,577 |
| **Father** | | | | | | |
| Sample: *Child age 4 months* | 2,561 | 1,373 | 0,003 | 0,031 | 0,930 | 0,934 |
| Sample: *Child age 8 months* | 3,969 | 2,764 | 0,002 | 0,016 | 0,927 | 0,930 |
| Sample: *Child age 24 months* | 4,215 | 4,313 | 0,001 | 0,002 | 0,845 | 0,880 |

Source: Spanish Labor Force Survey (EPA).
Note: Pre-reform observations correspond to individuals with a child born within 6 months before the reform (March 2007) and Post-reform observations correspond to individuals with a child born within 6 months after the reform. The raw Sample: *"child age 4 months"* displays the estimated results for individuals interviewed in the third or fourth quarter in 2006 or in any quarter in 2007. The raw Sample *"child age 8 months"* extends the sample by including also individuals interviewed in the first and second quarter in 2008. The raw Sample *"child age 24 months"* includes interviews conducted in the last two quarters in 2008 and in 2009.

B. Fertility - Labor Force Survey (EPA)

| | N.obs. Pre-reform | N.obs. Post-reform | Additional birth(s) Pre-reform | Additional birth(s) Post-reform |
|---|---|---|---|---|
| Sample : *All mothers* | 9,605 | 9,584 | 0,130 | 0,076 |

Source: Spanish Labor Force Survey (EPA).
Note: Pre-reform observations correspond to individuals with a child born within 6 months before the reform (March 2007) and Post-reform observations correspond to individuals with a child born within 6 months after the reform. The sample includes the fertility records for the period 2008 to 2010 of all women who had children around the time of the reform.

Table 2 (Cont) : Descriptive statistics

C. Labor market outcomes - Social Security Records

| | Earnings first 6 months Pre-reform | Earnings first 6 months Post-reform | Earnings first 12 months Pre-reform | Earnings first 12 months Post-reform | Employed after 6 months Pre-reform | Employed after 6 months Post-reform | Unpaid parental leave Pre-reform | Unpaid parental leave Post-reform |
|---|---|---|---|---|---|---|---|---|
| **Mother** | 5,861 | 6,191 | 11,951 | 12,681 | 0,602 | 0,640 | 0,073 | 0,055 |
| N.obs. | 6,608 | 6,921 | 6,608 | 6,921 | 6,615 | 6,926 | 6,615 | 6,926 |
| **Father** | 8,642 | 8,723 | 17,590 | 17,587 | 0,869 | 0,858 | 0,002 | 0,002 |
| N.obs. | 7,461 | 7,885 | 7,461 | 7,885 | 7,466 | 7,891 | 7,466 | 7,891 |

Source: Social Security's 2009 continuous sample of working histories.
Note: Pre-reform observations correspond to individuals with a child born within 6 months before the reform (March 2007) and Post-reform observations correspond to individuals with a child born within 6 months after the reform.

Table 3: Balance in covariates test

| | Age | Primary education | High School degree | College | Immigrant |
|---|---|---|---|---|---|
| **Mothers** | | | | | |
| Sample : *Child age 4 months* | -0.071 | -0.021 | -0.023 | -0.008 | -0.049** |
| | [0.296] | [0.016] | [0.017] | [0.016] | [0.023] |
| N.obs. | 3,935 | 3,935 | 3,935 | 3,935 | 3,935 |
| Sample: *Child age 8 months* | 0.099 | -0.014 | -0.007 | -0.007 | -0.048*** |
| | [0.229] | [0.012] | [0.013] | [0.013] | [0.017] |
| N.obs. | 6,734 | 6,734 | 6,734 | 6,734 | 6,734 |
| Sample: *Child age 24 months* | 0.039 | -0.022** | -0.018 | -0.024** | -0.029* |
| | [0.202] | [0.011] | [0.011] | [0.011] | [0.015] |
| N.obs. | 8,528 | 8,528 | 8,528 | 8,528 | 8,528 |
| **Father** | | | | | |
| Sample: *Child age 4 months* | 0.511 | 0.039** | 0.020 | 0.039** | -0.000* |
| | [0.335] | [0.019] | [0.020] | [0.018] | [0.000] |
| N.obs. | 3,935 | 3,935 | 3,935 | 3,935 | 3,935 |
| Sample: *Child age 8 months* | 0.549** | 0.025 | 0.019 | 0.023* | 0.000 |
| | [0.257] | [0.015] | [0.016] | [0.014] | [0.000] |
| N.obs. | 6,734 | 6,734 | 6,734 | 6,734 | 6,734 |
| Sample: *Child age 24 months* | 0.195 | 0.020 | 0.016 | -0.001 | 0.000 |
| | [0.221] | [0.013] | [0.013] | [0.012] | [0.000] |
| N.obs. | 8,528 | 8,528 | 8,528 | 8,528 | 8,528 |

Data Source: Spanish Labor Force Survey (EPA).
Note: Robust standard errors in brackets. The reported coefficients are for the binary indicator taking value 1 for months after March 2007. The estimates are obtained by comparing the characteristics of individuals with children born 6 months before the reform (March 2007) and 6 months after. The raw Sample: *"child age 4 months"* displays the estimated results for individuals interviewed in the third or fourth quarter in 2006 or in any quarter in 2007. The raw Sample *"child age 8 months"* extends the sample by including also individuals interviewed in the first and second quarter in 2008. The raw Sample *"child age 24 months"* includes interviews conducted in the last two quarters in 2008 and in 2009.
*** p<0.01, **p<0.05, *p<0.1.

Table 4. Validity check: Bunching of births at the threshold

| Days of birth around the reform | +/- 7 days | +/- 14 days | +/- 21 days | +/- 42 days | +/- 56 days | +/- 77 days |
|---|---|---|---|---|---|---|
| Daily n. of births | 42 | -109 | -81 | -48 | -36 | -1 |
| | [91] | [107] | [92] | [67] | [57] | [49] |
| Log n. of births | 0.032 | -0.091 | -0.068 | -0.043 | -0.033 | -0.005 |
| | [0.074] | [0.087] | [0.075] | [0.055] | [0.047] | [0.040] |
| Linear trends | NO | YES | YES | YES | YES | YES |
| Quadratic trends | NO | NO | NO | NO | NO | NO |
| N.obs. | 14 | 28 | 42 | 84 | 112 | 154 |

Source: Birth-certificate data.

Note: Robust standard errors in brackets. The reported coefficients are for the binary indicator taking value 1 for months after March 2007. The sample includes all days in the specified window around March 24, 2007. The outcome variable is the (log) daily number of births. The main explanatory variable is an indicator for birthdates on or after March 24, 2007. In all but the first column, we control for a linear trend in date of birth (the running variable, centered at 0 in March 24, 2007), interacted with the main explanatory variable. One asterisk indicates significance at 90%.

Table 5: Effects on the leave-taking behavior and employment of fathers

| | On parental leave (1) | On parental leave (2) | On parental leave (3) | On leave other family reasons (4) | On leave other family reasons (5) | On leave other family reasons (6) | Employed (7) | Employed (8) | Employed (9) |
|---|---|---|---|---|---|---|---|---|---|
| Months of birth around the reform | +/-6 | +/-3 | +/-9 | +/-6 | +/-3 | +/-9 | +/-6 | +/-3 | +/-9 |
| Sample: | 0.013* | 0.019*** | 0.017** | 0.002 | 0.005* | 0.002 | 0.014 | 0.012 | 0.019 |
| *Child age 4 months* | [0.007] | [0.005] | [0.008] | [0.003] | [0.003] | [0.003] | [0.015] | [0.011] | [0.017] |
| N.obs. | 3,934 | 1,961 | 6,038 | 3,934 | 1,961 | 6,038 | 3,934 | 1,961 | 6,038 |
| Sample: | 0.009** | 0.010*** | 0.009** | 0.002 | 0.003* | 0.001 | 0.002 | 0.002 | 0.001 |
| *Child age 8 months* | [0.004] | [0.003] | [0.004] | [0.002] | [0.002] | [0.002] | [0.002] | [0.009] | [0.014] |
| N.obs. | 6,733 | 3,345 | 10,223 | 6,733 | 3,345 | 10,223 | 6,733 | 3,345 | 10,223 |
| Sample: | 0.004** | 0.002 | 0.004 | 0.001 | 0.000 | 0.001 | 0.001 | 0.010 | 0.012 |
| *Child age 24 months* | [0.002] | [0.002] | [0.002] | [0.001] | [0.000] | [0.001] | [0.001] | [0.010] | [0.018] |
| N.obs. | 8,528 | 4,130 | 12,967 | 8,528 | 4,130 | 12,967 | 8,528 | 4,130 | 12,967 |
| Liner trend in m | YES | NO | YES | YES | NO | YES | YES | NO | YES |
| Quadratic trend in m | NO | NO | YES | NO | NO | YES | NO | NO | YES |

Data Source: Spanish Labor Force Survey (EPA).
Note: Robust standard errors in brackets. The reported coefficients are for the binary indicator taking value 1 for months after March 2007. Individual controls include: a third order polynomial on the age of the father, dummies for the level of education (less than Primary education, Primary education, High School degree or College) and an immigrant indicator. In columns (1), (4) and (7) the outcomes of fathers with children born 6 months before the reform (March 2007) are compared to those of children born 6 months after. Columns (2), (5) and (84) compare the outcomes of fathers with children born 3 months around the reform and columns (3), (6) and (9) those with children born 9 months around the reform. The raw Sample: *"child age 4 months"* displays the estimated results for individuals interviewed in the third or fourth quarter in 2006 or in any quarter in 2007. The raw Sample *"child age 8 months"* extends the sample by including also individuals interviewed in the first and second quarter in 2008. The raw Sample *"child age 24 months"* includes interviews conducted in the last two quarters in 2008 and in 2009.
*** p<0.01, **p<0.05, *p<0.1.

Table 6: Difference in difference estimation of the effects of the reform on the leave-taking behavior and labor market outcomes

| | On parental leave (1) | Employed (2) | On parental leave (3) | Employed (4) |
|---|---|---|---|---|
| | Father | Father | Mother | Mother |
| Months of birth around the reform | +/-18 | +/-18 | +/-18 | +/-18 |
| Sample DiD1 | 0.014*** | -0.015 | 0.018 | 0.071*** |
| Q305 to Q408 | [0.004] | [0.013] | [0.015] | [0.024] |
| N.obs. | 28,703 | 28,703 | 28,709 | 28,709 |
| Sample DiD2 | 0.008*** | 0.007 | 0.012 | 0.019 |
| Q305 to Q409 | [0.003] | [0.012] | [0.010] | [0.018] |
| N.obs. | 46,241 | 46,241 | 46,251 | 46,251 |
| Sample DiD3 | 0.005** | 0.014 | 0.009 | 0.012 |
| Q305 to Q410 | [0.002] | [0.010] | [0.007] | [0.015] |
| N.obs. | 64,138 | 64,138 | 64,148 | 64,148 |
| Linear trend in m | YES | YES | YES | YES |
| Quadratic in m | YES | YES | YES | YES |
| month fixed effects | YES | YES | YES | YES |

Source: Spanish Labor Force Survey (EPA).
Note: Robust standard errors in brackets. The reported coefficients are for the binary indicator taking value 1 for months after March 2007 when "seasonality" is accounted for by including month of birth fixed effects. Individual controls include: a third order polynomial on age, dummies for the level of education (less than Primary education, Primary education, High School degree or College) and an immigrant indicator. In all columns the outcomes of individuals with children born 18 months before the reform (from October 2005 to March 2007) are compared to those of children born 18 months after (from April 2007 to September 2008). The raw Sample DD1 displays the estimated results for individuals interviewed in the third or fourth quarter in 2005 or in any quarter in 2006, 2007 and 2008. The raw Sample DD2 extends DD1 by including individuals interviewed in 2009, and the raw Sample DD3 adds interviews in 2010. *** p<0.01, **p<0.05, *p<0.1.

Table 7: Effects on the earnings, employment and parental leave of fathers

| Months of birth around the reform | +/-3 | +/-6 | +/-9 | +/-9 |
|---|---|---|---|---|
| Earnings first 6 months | -7 | -129 | -353*** | 58 |
| | [102] | [137] | [128] | [172] |
| Earnings first 12 months | 117 | -175 | -467* | 87 |
| | [205] | [273] | [258] | [341] |
| Employed after 6 months | 0.008 | 0.002 | 0.003 | -0.002 |
| | [0.123] | [0.010] | [0.013] | [0.007] |
| Unpaid Parental leave | 0.001 | 0.001 | -0.001 | 0.001 |
| | [0.001] | [0.002] | [0.001] | [0.002] |
| N.obs. | 7406 | 15346 | 22834 | 22834 |
| Linear trend in m | NO | YES | YES | YES |
| Quadratic in m | NO | NO | NO | YES |

Data source: Social Security's 2009 continuous sample of working histories.
Note: Robust standard errors in parentheses. The reported coefficients are for the binary indicator taking value 1 for months after March 2007. The dependent variables are: gross earnings of the father during the first 6 or 12 months after birth, an employment dummy measured 6 months after birth, and an indicator for the job held on the month of birth being suspended at some point due to parental leave. The coefficients are those for the dummy indicating children born after March 2007. Individual controls include: a third order polynomial for the age of the father, dummies for the level of education, and indicators for whether the father had a permanent job or was self-employed three months before birth. Each column uses a sample that includes a different number of months of birth (3, 6 or 9) around the threshold. *** p<0.01, **p<0.05, *p<0.1.

Table 8: Effects on the leave-taking behavior and employment of mothers

| | On parental leave (1) | On parental leave (2) | On parental leave (3) | On leave other family reasons (4) | On leave other family reasons (5) | On leave other family reasons (6) | Employed (7) | Employed (8) | Employed (9) |
|---|---|---|---|---|---|---|---|---|---|
| Month of birth around the reform | +/-6 | +/-3 | +/-9 | +/-6 | +/-3 | +/-9 | +/-6 | +/-3 | +/-9 |
| Sample: | 0.024 | 0.013 | 0.015 | -0.010 | -0.023 | -0.010 | 0.061** | 0.059 | 0.080** |
| *Child age 4 months* | [0.026] | [0.037] | [0.032] | [0.015] | [0.023] | [0.019] | [0.029] | [0.043] | [0.037] |
| N.obs. | 3,935 | 1,961 | 6,040 | 3,935 | 1,961 | 6,040 | 3,935 | 1,961 | 6,04 |
| Sample: | 0.011 | 0.003 | 0.007 | -0.002 | -0.017 | -0.010 | 0.060*** | 0.099*** | 0.086*** |
| *Child age 8 months* | [0.016] | [0.024] | [0.021] | [0.011] | [0.017] | [0.014] | [0.023] | [0.034] | [0.029] |
| N.obs. | 6,734 | 1,961 | 10,225 | 6,734 | 1,961 | 10,225 | 6,734 | 3,345 | 10,225 |
| Sample: | 0.000 | -0.005 | 0.001 | 0.014** | 0.005 | 0.009 | -0.014 | -0.046 | -0.038 |
| *Child age 24 months* | [0.005] | [0.008] | [0.007] | [0.006] | [0.008] | [0.008] | [0.020] | [0.030] | [0.026] |
| N.obs. | 8,528 | 4,130 | 12,967 | 8,528 | 4,130 | 12,967 | 8,528 | 4,130 | 12,967 |
| Liner trend in m | YES | YES | YES | YES | YES | YES | YES | YES | YES |
| Quadratic trend in m | NO | NO | YES | NO | NO | YES | NO | NO | YES |

Data Source: Spanish Labor Force Survey (EPA).

Note: Robust standard errors in parenthesis. The reported coefficients are for the binary indicator taking value 1 for months after March 2007. Individual controls include: a third order polynomial on the age of the mother, dummies for the level of education (less than Primary education, Primary education, High School degree or College) and an immigrant indicator. In columns (1), (4) and (7) the outcomes of mothers with children born 6 months before the reform (March 2007) are compared to those of children born 6 months after. Columns (2), (5) and (8) compare the outcomes of mothers with children born 3 months around the reform and columns (3), (6) and (9) those with children born 9 months around the reform. The raw Sample: *"child age 4 months"* displays the estimated results for individuals interviewed in the third or fourth quarter in 2006 or in any quarter in 2007. The raw Sample *"child age 8 months"* extends the sample by including also individuals interviewed in the first and second quarter in 2008. The raw Sample *"child age 24 months"* includes interviews conducted in the last two quarters in 2008 and in 2009. *** p<0.01, **p<0.05, *p<0.1.

Table 9: Effects on the earnings, employment and parental leave of mothers

| Months of birth around the reform | +/-3 | +/-6 | +/-9 | +/-9 |
|---|---|---|---|---|
| Earnings first 6 months | 276*** | 84 | 160 | 96 |
| | [91] | [132] | [106] | [169] |
| Earnings first 12 months | 701*** | 346 | 361* | 324 |
| | [187] | [269] | [217] | [345] |
| Employed after 6 months | 0.003 | 0.018 | 0.013 | 0.033*** |
| | [0.019] | [0.015] | [0.019] | [0.010] |
| Unpaid Parental leave | -0.020*** | -0.019** | -0.025*** | -0.013 |
| | [0.006] | [0.009] | [0.007] | [0.011] |
| N.obs. | 6,579 | 13,541 | 20,313 | 20,313 |
| Linear trend in m | NO | YES | YES | YES |
| Quadratic in m | NO | NO | NO | YES |

Data source: Social Security's 2009 continuous sample of working histories.
Note: Robust standard errors in brackets. The reported coefficients are for the binary indicator taking value 1 for months after March 2007. The dependent variables are: gross earnings of the father during the first 6 or 12 months after birth, an employment dummy measured 6 months after birth, and an indicator for the job held on the month of birth being suspended at some point due to parental leave. The coefficients are those for the dummy indicating children born after March 2007. Individual controls include: a third order polynomial for the age of the father, dummies for the level of education, and indicators for whether the father had a permanent job or was self-employed three months before birth. Each column uses a sample that includes a different number of months of birth (3, 6 or 9) around the threshold. *** p<0.01, **p<0.05, *p<0.1.

Table 10: Heterogeneous effect on leave-taking behavior and employment by education

| | On parental leave (1) | Employed (2) | On parental leave (3) | Employed (4) |
|---|---|---|---|---|
| | Father | Father | Mother | Mother |
| *College Graduates* | 0.006 | 0.011 | 0.018 | 0.012 |
| | [0.008] | [0.014] | [0.033] | [0.036] |
| N.obs. | 1,393 | 1,393 | 2,032 | 2,032 |
| | | | | |
| *Non College Graduates* | 0.010** | 0.002 | 0.005 | 0.081*** |
| | [0.004] | [0.015] | [0.018] | [0.030] |
| N.obs. | 5,340 | 5,340 | 4,702 | 4,702 |
| | | | | |
| Linear trend in m | YES | YES | YES | YES |

Source: Spanish Labor Force Survey (EPA).

Note: Robust standard errors in parenthesis. The reported coefficients are for the binary indicator taking value 1 for months after March 2007. Individual controls include: a third order polynomial on the age of the father or mother and an immigrant indicator. In all columns the outcomes of mothers with children born 6 months before the reform (March 2007) are compared to those of children born 6 months after. The results are obtained from the sample *"child age 8 months"* which includes interviews in the third or fourth quarter in 2006, in any quarter in 2007 or the first two quarters in 2008. *** p<0.01, **p<0.05, *p<0.1.

Table 11: Effects on fertility (Labor Force Survey)

| | N.obs. | Additional birth(s) |
|---|---|---|
| *All mothers* | 19,189 | -0.019** |
| | | [0.009] |
| *College Graduates* | 5,753 | -0.002 |
| | | [0.016] |
| *Non College Graduates* | 13,436 | -0.024** |
| | | [0.010] |
| *Younger than 35* | 10,374 | -0.012 |
| | | [0.012] |
| *Older than 35* | 8,815 | -0.027** |
| | | [0.012] |

Source: Spanish Labor Force Survey (EPA).
Note: Robust standard errors in parenthesis. The reported coefficients are for the binary indicator taking value 1 for months after March 2007. Individual controls include: a third order polynomial on the age of the mother and an immigrant indicator. The estimates in all columns are obtained from comparing the outcomes of mothers with children born within 6 months before the reform (March 2007) to those of children born within 6 months after. The sample includes observations for the years 2008, 2009 and 2010. *** p<0.01, **p<0.05, *p<0.1.

Table 12. Effects on birth spacing, number of days (Birth-certificate data)

| Days of birth around the reform | +/- 7 days | +/- 14 days | +/- 21 days | +/- 42 days | +/- 56 days | +/- 77 days |
|---|---|---|---|---|---|---|
| *Number of days between births* | 13.5 | 26.6 | 26.7* | 17.9* | 21.7** | 16.8** |
| | (26.1) | (18.3) | (14.8) | (10.5) | (9.1) | (7.7) |
| Mean | 1,251 | 1,251 | 1,249 | 1,250 | 1,250 | 1,249 |
| Effect as % of mean | 1.1% | 2.1% | 2.1% | 1.4% | 1.7% | 1.3% |
| N.obs. | 6,473 | 12,870 | 19,246 | 38,523 | 51,628 | 71,341 |

Source: Birth-certificate data.
Note: Robust standard errors in parenthesis. The reported coefficients are for the binary indicator taking value 1 for months after March 2007. The sample includes all mothers who had another child after the reform within 6 years. For those women the first child was born in a certain window of days around March 24, 2007. The outcome variable is the number of days between the birth that took place around March 2007 and the next birth. In all columns, we control for a linear trend in date of birth (the running variable, centered at 0 in March 24, 2007), interacted with the post reform indicator. *** p<0.01, **p<0.05, *p<0.1.

Table 13. Effects on fertility within 2 years after (previous) birth (Birth-certificate data)

| | | | | | | |
|---|---|---|---|---|---|---|
| | Older mothers (>31) | | | | | |
| Day of birth around the reform | +/- 7 days | +/- 14 days | +/- 21 days | +/- 42 days | +/- 56 days | +/- 77 days |
| Paternity | -0.0067 | -0.0057 | -0.0085 | -0.0074** | -0.0060* | -0.0069** |
| leave | (0.0046) | (0.004) | (0.0052) | (0.0038) | (0.0033) | (0.0028) |
| Average | 0.053 | 0.053 | 0.054 | 0.055 | 0.055 | 0.054 |
| Coeff./average | -12.6% | -10.7% | -15.8% | -13.4% | -10.9% | -12.7% |
| N. obs. | 9,677 | 19,057 | 28,525 | 57,034 | 76,476 | 105,700 |
| | Younger mothers (<32) | | | | | |
| Paternity | -0.0071 | -0.0061 | -0.0073 | -0.0056 | -0.0043 | -0.0030 |
| leave | (0.0060) | (0.0084) | (0.0069) | (0.0049) | (0.0043) | (0.0037) |
| Average | 0.080 | 0.081 | 0.082 | 0.083 | 0.083 | 0.083 |
| Coeff./average | -8.9% | -7.5% | -8.9% | -0.7% | -5.2% | -3.6% |
| N. obs. | 8,143 | 16,096 | 24,062 | 48,100 | 64,582 | 89,590 |

Source: Birth-certificate data.
Note: Robust standard errors in parenthesis. The reported coefficients are for the binary indicator taking value 1 for months after March 2007. The sample includes all mothers that gave birth in a certain window of days around March 24, 2007. The outcome variable is an indicator for the mother having another child within the following 2 years. The main explanatory variable is an indicator for the (initial) birth taking place on or after March 24, 2007. In all but the first column, we control for a linear trend in date of birth (the running variable, centered at 0 in March 24, 2007), interacted with the main explanatory variable. *** p<0.01, **p<0.05, *p<0.1.

## Appendix

Table A1: Effects on joint leave-taking behavior of parents

| | Only_mother (1) | Only_father (2) | Both (3) | Only_mother (4) | Only_father (5) | Both (6) |
|---|---|---|---|---|---|---|
| Months of birth around the reform | [+/-]6 | [+/-]6 | [+/-]6 | [+/-]9 | [+/-]9 | [+/-]9 |
| Sample: | 0.025 | 0.008 | 0.005 | 0.016 | 0.011* | 0.006 |
| *Child age 4 months* | [0.026] | [0.005] | [0.004] | [0.031] | [0.006] | [0.005] |
| N.obs. | 3,934 | 3,934 | 3,934 | 6,038 | 6,038 | 6,038 |
| | | | | | | |
| Sample: | 0.010 | 0.006** | 0.003 | 0.006 | 0.006* | 0.003 |
| *Child age 8 months* | [0.016] | [0.003] | [0.003] | [0.020] | [0.003] | [0.003] |
| N.obs. | 6,733 | 6,733 | 6,733 | 10,223 | 10,223 | 10,223 |
| | | | | | | |
| Sample: | -0.002 | 0.002 | 0.003* | -0.001 | 0.002 | 0.002 |
| *Child age 24 months* | [0.005] | [0.002] | [0.002] | [0.007] | [0.003] | [0.002] |
| N.obs. | 8,528 | 5,696 | 5,696 | 12,967 | 8,736 | 8,736 |
| | | | | | | |
| Liner trend in m | YES | YES | YES | YES | YES | YES |
| Quadratic trend in m | NO | NO | NO | YES | YES | YES |

Source: Spanish Labor Force Survey (EPA).

Note: Robust standard errors in brackets. The estimates correspond to the coefficient on Post in equation (1). The set of covariates is extended to include the age, education and immigrant condition of both spouses. In column (1) and (4) the depend variable takes value 1 if only the mother is on leave. In column (2) and (5) the dependent variable takes value 1 if only the father is on leave, and in column (3) and (6) the dependent variable takes value 1 when both spouses are on leave. The raw Sample: *"child age 4 months"* displays the estimated results for individuals interviewed in the third or fourth quarter in 2006 or in any quarter in 2007. The raw Sample *"child age 8 months"* extends the sample by including also individuals interviewed in the first and second quarter in 2008. The raw Sample *"child age 24 months"* includes interviews conducted in the last two quarters in 2008 and in 2009. *** p<0.01, **p<0.05, *p<0.1.

Table A2. Effects on fertility within 4 years after (previous) birth (Birth-certificate data)

| | | | Older mothers (>31) | | | |
|---|---|---|---|---|---|---|
| Day of birth around the reform | +/- 7 days | +/- 14 days | +/- 21 days | +/- 42 days | +/- 56 days | +/- 77 days |
| Paternity leave | -0.0068 | -0.0026 | -0.0078 | -0.0053 | -0.0075 | -0.0116** |
| | (0.0080) | (0.0115) | (0.0093) | (0.0066) | (0.0057) | (0.0049) |
| N. obs. | 9,677 | 19,057 | 28,525 | 57,034 | 76,476 | 105,700 |
| | | | Younger mothers (<32) | | | |
| Paternity leave | -0.0062 | -0.0137 | -0.0096 | -0.0026 | 0.0028 | -0.0000 |
| | (0.0102) | (0.0145) | (0.0118) | (0.0084) | (0.0072) | (0.0062) |
| N. obs. | 8,143 | 16,096 | 24,062 | 48,100 | 64,582 | 89,590 |

Source: Birth-certificate data.
Note: Robust standard errors in parenthesis. The reported coefficients are for the binary indicator taking value 1 for months after March 2007. The sample includes all mothers that gave birth in a certain window of days around March 24, 2007. The outcome variable is an indicator for the mother having another child within the following 4 years. The main explanatory variable is an indicator for the (initial) birth taking place on or after March 24, 2007. In all but the first column, we control for a linear trend in date of birth (the running variable, centered at 0 in March 24, 2007), interacted with the main explanatory variable. *** p<0.01, **p<0.05, *p<0.1.