Does Paternity Leave Reduce Fertility?*

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Abstract: We find that the introduction of two weeks of paid paternity leave in Spain in 2007 led to delays in subsequent fertility. Following a regression discontinuity design and using rich administrative data, we show that parents who were (just) entitled to the new paternity leave took longer to have another child compared to (just) ineligible parents. We also show that older eligible couples were less likely to have an additional child within the following six years after the introduction of the reform. We provide evidence in support of two potentially complementary channels behind the negative effects on subsequent fertility. First, fathers' increasing involvement in childcare led to higher labor force attachment among mothers. This may have raised the opportunity cost of an additional child. We also find that men reported lower desired fertility after the reform, possibly due to their increased awareness of the costs of childrearing, or to a shift in preferences from child quantity to quality.

Keywords: Paternity leave, fertility, labor market, gender, natural experiment.

JEL codes: J48, J13, J16.

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1. Introduction

We show that the introduction of two weeks of paid paternity leave in Spain in 2007 led to an increase in birth spacing, which may have led to fewer subsequent births among older couples. Following a regression discontinuity design, we find that parents who were (just) entitled to the new paternity leave when they had a child in 2007 took longer to have another child, compared with (just) ineligible parents. We also show that eligible couples were less likely to have an additional child within the following six years after the introduction of the reform. This drop in subsequent fertility is mostly driven by women over 30. A persistent increase in fathers' involvement in childcare, combined with a higher employment rate of mothers the first year after birth, and a fall in the desired fertility of men, are potential candidates to explain the fertility delay.

Essentially all countries in the world provide some form of paid maternity leave, while many also offer paternity leave (Rossin-Slater 2018). On average, public expenditure on maternity and parental leave was $12,300 per child born in OECD countries in 2013 (OECD Family database). There is some evidence that increases in the generosity of maternity leave can boost fertility (Olivetti and Petrongolo 2017).[1] Lalive and Zweimüller (2009) found that an extension in the duration of maternity leave in Austria in 1990 led to a substantial increase in fertility, and Raute (2017) found that financial incentives for high-earnings women to take maternity leave also led to fertility gains. On the other hand, Dahl et al. (2016) find no fertility effects of two extensions of paid maternity leave in Norway.

Parental leave entitlements reserved to fathers and non-transferable to mothers ("use-it-or-lose-it") seek to increase the participation of fathers in childcare activities. Despite

[1] A rich recent literature, surveyed in Olivetti and Petrongolo (2017) and Rossin-Slater (2018), has also documented the effects of maternity leave (extensions) on mothers' careers and child outcomes.

the substantial improvement in female labor market opportunities in recent decades, women continue to devote much more time to unpaid and care work than men. A more balanced distribution of unpaid work within households may allow women to spend more time and effort in paid work, fostering their professional careers. A greater involvement of fathers in childrearing may also affect employers' decisions regarding the hiring and promotion of women, potentially reducing gender disparities in the labor market.

Reserving parental leave time for fathers may also affect households' fertility decisions. Increasing the duration and/or generosity of parental leave may encourage fertility by providing more options for parents to balance work and family. With the increased labor market involvement of women, the distribution of the childcare burden between mothers and fathers has become an important determinant of fertility (Feyrer et al. 2008, Doepke and Kindermann 2016). The introduction of paternity leave permits, by potentially altering the allocation of child care duties between spouses, may affect their desired number of children and fertility outcomes.

There is limited evidence to date on the effects of leave provisions for fathers, due to their more recent introduction (Olivetti and Petrongolo 2017). A small number of papers have documented large effects on take-up in the US, Norway, Sweden and Canada (Bartel et al. 2018, Cools et al. 2015, Dahl et al. 2014, Ekberg et al. 2013, Patnaik 2016).

Descriptively, fathers who take more parental leave are more involved in childcare activities later on (Almquist and Duvander 2014, Bünning 2015, Huerta et al. 2013, Nepomnyaschy and Waldfogel 2007, Tanaka and Waldfogel 2007) and work fewer hours (Bünning and Pollmann-Schult 2016). The evidence is more mixed in studies that aim at identifying the causal effect of paternity leave on men and women's labor supply, childcare involvement, and earnings (Cools et al. 2015, Dahl et al. 2014, Dunatchik and

Özcan 2017, Ekberg et al. 2013, Kluve and Tamm 2013, Patnaik 2016, Rege and Solli 2013, Tamm 2018).

Recent work has also reported that an increase in fathers' share of parental leave increased marital separation rates in Sweden (Avdic and Karimi 2018), although Dahl et al. (2014) and Cools et al. (2015) found no effect of paternity leave on marital stability in Norway. Few studies have analyzed the effect of paternity leave on fertility, and those that have, report zero effects. Cools et al. (2015), Dahl et al. (2014), and Kotsadam and Finseraas (2001) found no effects of paternity leave extensions on fertility in Norway, and Bartel et al. (2018) reported similar results for the US.

To explore the effects of the introduction of two weeks of paternity leave in Spain, we use administrative data and focus on families who had a child just before and after the reform in March 2007. First we show that take-up was very high among new fathers (Figure 1).[2] The analysis of birth-certificate data shows that the reform led couples to delay subsequent births (see Figure 2), resulting in fewer births in the following years, especially among older couples (see Figure 3). Using Social Security data, we find no effects on labor market outcomes for fathers, but mothers in eligible families had higher employment rates 6 months after childbirth (see Figure 4). We also find that fathers in eligible households reported spending more time on childcare, and this effect was persistent.

We provide evidence in support of two (potentially complementary) channels driving the negative fertility effects. First, fathers’ increasing involvement in childcare may have improved mothers’ labor force attachment, as reflected in their higher employment rates after birth, which may have increased the opportunity cost of an additional child. We also

[2] Escot et al. (2014) analyze take-up effects of the Spanish reform with Labor Force Survey data.

find that men reported lower desired fertility after the reform (see Figure 5), which may have resulted from the paternity leave period raising their awareness about the full cost of having children. Alternatively, spending more time with their child may have shifted their preferences in favor of child quality (versus quantity).

The fertility effects that we find may generalize to other Southern and Eastern European countries, where women still shoulder the bulk of home production. Before 2007, fathers in Spain were making essentially no use of parental leave (Figure 1). Mothers' employment rates were low (Figure 6), and women were doing most of the childcare and housework (Figure 7). According to the Spanish Time Use Survey, in 2002-03 women spent 4.2 hours a day in housework and childcare, compared with 1.3 for men. At the same time, men had higher desired fertility than women (Figure 8), which was not the case in Northern European countries.

These features may have made the introduction of paternity leave more effective in terms of increasing fathers' childcare time and women's attachment to the labor force, perhaps with the side effect of lowering the desired fertility of men relative to women. These effects may be more muted in countries where both market and household work were less unequally distributed between mothers and fathers, at the time of the paternity leave extensions.

We acknowledge that we only provide one data point regarding potential negative effects of paternity leave on fertility. The effects are modest in size, and precision is limited, so that we need more research in order to conclude that paternity leave can lead to lower fertility in other contexts.

The remainder of the paper is organized as follows. We provide a brief description of the institutional setting in Section 2. We then describe the methodology and the data sources (Sections 3 and 4), explain the results in Section 5, and conclude in Section 6.

2. Institutional setting

Since 1999, Spain granted 6 weeks of compulsory maternity leave (at full pay), plus 2 days of paid job absence for fathers. In addition, families were granted 10 weeks of parental leave, also at full pay, which could be taken by mothers or fathers, or shared between them (see Appendix Table 1). As shown in Figure 1, very few fathers used any parental leave. After the paid leave period, either parent could take unpaid leave for up to 3 years, with a right to return to the same job. One of the parents could also reduce working hours up to 50% (with a proportional reduction in pay) until the child turned 6. In practice, very few fathers made use of either the unpaid leave or the reduction in hours. In the 2006 Spanish Labor Force Survey, 3.8% of women with children up to 3 years of age reported being on unpaid leave, compared with less than 0.1% of men. The percentage of women with children younger than 6 that reported working part-time due to family responsibilities was 17%, while it was only 0.3% among men.

In 2007, the national government introduced a new 13-day paternity leave period, fully compensated, which could be taken by fathers either at the same time or immediately after the maternity leave period.[3] New fathers were eligible starting from births taking place on March 24, 2007 (the day after the law was published), provided they were affiliated to Social Security and had worked for at least 180 days over the previous 7 years. As shown in Figure 1, take-up was high, with about 55% of new fathers using it in 2008. The paternity leave permit was later extended to four weeks in January 2017, and to five in July 2018.

[3] The introduction of the paternity leave period was regulated by Law 3/2007, passed on March 22, 2007 and published on March 23, 2007.

3. Methodology

We analyze the effect of the introduction of two weeks of paternity leave on fertility outcomes. We follow a regression discontinuity design, based on the fact that the reform came into effect on March 24, 2007, such that the families who had a child from that date on were eligible for the 13 days of leave reserved for fathers, while those having a child before were not. Our running variable is thus the date of birth of the child. We restrict the sample to families who had a child within a few weeks around the threshold (which we call their "reference child"), and study their subsequent fertility decisions. We estimate intent-to-treat (ITT) effects, since we observe eligibility (date of birth) but not actual take-up at the individual level. Figure 1 suggests that between 50 and 60% of new fathers took up the new paternity leave.[4] Thus, our results suggest effects on the treated that are up to twice as large as our ITT estimates.

We estimate the following equation:

$$(1) \quad Y_{id} = \alpha + \beta T_{id} + \gamma_1 f(d) + \gamma_2 f(d) T_{id} + \delta X_{id} + \varepsilon_{id},$$

where *Y* is either the number of days between the birthdate *d* of the reference child and the birthdate of the next child to the same mother *i*, or an indicator for whether the mother had another child within 2, 4 or 6 years of the birth of the reference child. *T* is an indicator for the reference child having been born on or after March 24, 2007, and *f* is a polynomial in the running variable *d*. Date of birth *d* is normalized to 0 on March 24, 2007.

We include family and time-related controls *X* (a third-order polynomial in age of the mother, a set of indicators for maternal education, and day of the week fixed effects for the date of birth). We explore a set of different windows around the threshold (from 7 to 90

[4] Information on parental leave claims is only available at the annual level. In 2007, the number of paternity leave claims amounted to about 46% of all births after March 23. In 2008, the number of claims amounted to about 55% of all births. Thus, we estimate an initial take-up rate close to 50%.

days). We also report the results of an optimal bandwidth selector, as well as a "donut" specification that excludes the 7 days right before and after the cutoff date. We also vary the order of the polynomial in the running variable. Standard errors are clustered at the date of birth level.

The identifying assumption is that having a child right before or right after March 24, 2007 is as good as random. If this assumption holds, we expect to observe no bunching in the number of births in the days right after the threshold, as well as family and newborn characteristics balanced around the threshold, on average. We test for these implications in section 4.

Our RD approach compares the short and longer-term outcomes of individuals having children around March 24, 2007 (up to 7 to 90 days before and after in the fertility analysis). The Great Recession hit the Spanish economy in 2008. The unemployment rate was 8.6% in the last quarter of 2007, and climbed up to 13.8% by the end of 2008. The short length of the time windows defined around March 24, 2007 in our RD specifications ensures that individuals in our sample faced similar macroeconomic conditions at the time of childbirth, as well as in the following months and years. Hence, we can plausibly isolate the effect of the reform separately from that of other aggregate factors.

In order to address potential seasonality concerns, we also estimate the following difference-in-differences specification, where we include children born before and after March 24 of the treatment year (2007), as well as the years before and after the reform was implemented (2006 and 2008):

$$(2)\quad Y_{idt} = \alpha + \beta_1 T_{idt} + \beta_2 T_{idt} I_{2007} + \gamma f(\delta) + \theta X_{id} + \mu_t + \varepsilon_{id},$$

Our coefficient of interest is now β_2, on the interaction between "after March 24 births" (T) and the 2007 indicator (I_{2007}), where δ stands for day of the year, and we control for year fixed effects (μ_t). This specification controls for systematic differences in subsequent fertility across families having a child in different (even if close) days of the year.

We also study the effect of the paternity leave introduction on labor market outcomes for mothers and fathers. In order to do so, we estimate equations (1) and (2), using month instead of exact date of birth as the running variable, since the exact date is not available in the administrative labor market data. We explore windows of 3, 6, and 9 months around the threshold. As main labor market outcomes, we study employment, earnings, unpaid parental leave-taking, and part-time work. We construct each of these variables at months 6, 12 and 24 after the month of birth of the reference child.

4. Data

The fertility analysis is conducted using birth-certificate data for years 2005-2013, from the Spanish National Statistical Institute. This data set includes information on the universe of births registered in Spain annually, and contains the month of birth of each child (and the previous child to the same mother), as well as measures of newborn health and family demographics. We supplement the publicly available microdata files with the exact date of birth of each child, as well as the birthdate of the previous child to the same mother. These two additional variables were purchased from the National Statistical Institute.

In order to analyze effects on birth spacing, we select the sample of women who had a child by the end of our sample period (Deceber 31, 2013), and whose previous child was born in the neighborhood of March 24, 2007. We then calculate birth spacing as the number of days in between the two birthdates. To analyze effects on subsequent fertility, we keep all women who had a child in the neighborhood of March 24, 2007, and construct individual-level indicators for having another child within 2, 4 and 6 years after the date of birth of the reference child.

We use register data from Social Security records to analyze the effects of the paternity leave introduction on labor market outcomes of fathers and mothers. The data are publicly available and provide complete work histories for a 4% representative sample of all individuals affiliated with Social Security in a given year (including the unemployed). We merge the samples for years 2011 to 2015.[5] Since the data are longitudinal, we can construct work histories for adults residing with a child born on the relevant months around the threshold. We construct indicators for employment, unpaid parental leave, and part-time work, 6, 12, and 24 months after the month of birth of the reference child. We also construct total earnings during the first 6, 12, and 24 months after birth. Since paternity leave eligibility could potentially affect labor market status (and thus Social Security affiliation), we use Labor Force Survey data to test that the reform did not affect labor market participation in 2011-15, and thus the likelihood that an individual appears in our Social Security data.

Finally, in order to explore the channels behind our fertility results, we also exploit the Spanish Time Use Survey of 2009-10 to analyze effects on childcare time, the 2001, 2006, and 2011 Eurobarometer Survey to study effects on desired fertility, and the Labor Force Survey for marital stability effects.

5. Results

5.1. Birth spacing and fertility

We start by conducting validity checks for our RD strategy with the birth-certificate data. First, we test for potential bunching in the number of births around the threshold, which would suggest that families were able to sort into the treatment or control group, thus

[5] In order to minimize selection due to parental separation, the information on date of birth of the children living in the household is taken from the earliest year (after 2007) when the individual appears in the sample.

invalidating our identifying assumption. This could be the case if, for instance, (some) families postponed the date of a programmed cesarean section or induction to become eligible for the new paternity leave period.

Appendix Figure 1 shows the daily and weekly number of births before and after the paternity leave introduction.[6] Visually, there is no evidence of bunching in number of births around the threshold. We test for sorting formally by estimating regressions of the form of equation (1), aggregated at the day of birth level, where the outcome variable is the number (or the log number) of children born on a given date. Results are reported in Appendix Table 2. We present the results of seven different specifications, which vary in the number of weeks around the threshold included in the sample (between one and thirteen). The coefficients are all positive but small, and statistically indistinguishable from zero. These results indicate that there was no significant discontinuous jump in the number of births at the threshold, so that the introduction of the paternity leave period did not lead families to manipulate the date of birth to become eligible (or to avoid becoming eligible) for the new program.

Second, we test whether the families having a child immediately before and after the cutoff date were balanced in their observable characteristics. We estimate equation (1), using a range of mother and newborn characteristics as the outcome. Appendix Table 3 (Panel A) reports the results of five different specifications, for twelve family characteristics, including newborn health measures (birth weight, prematurity, etc) as well as maternal demographics (age, education, labor market and immigrant status) and fathers' age and education. The coefficients are all small, and only three out of 60 are statistically different from zero at the 95% confidence level.

[6] There are fewer births during the weekend, likely due to fewer scheduled deliveries.

Figure 9 shows the results graphically for two of the variables: age of the mother, and child weight at birth. We present average values for each of the two variables, by week of birth (where week 0 corresponds to March 24-30, 2007). The mothers in the sample were on average 32 years old at the time of childbirth, with no obvious jump around the last week of March 2007. About 7% of the children were born below 2,500 grams, and again there is no evidence of a discontinuity at the threshold.

The two validity checks thus support the main identifying assumption of our RD approach. We then move on to our main analysis, for the two fertility-related outcome variables. We first estimate the effect of the two weeks of paternity leave on birth spacing (the time elapsing between the reference birth and the next birth to the same mother). The average birth spacing in our sample is about 1,250 days, or about three and a half years (see Figure 2).

The main regression results are shown in Table 1. Panel A displays the results for seven different specifications, for the full sample of families. The results indicate that parents who were eligible for the two weeks of paternity leave took between 16 and 38 days longer to have their next child, a delay of between 1.3 and 3.1% (with respect to the mean birth interval of about 1,250 days, as shown in the table). Since these are ITT effects, and since we estimate a take-up rate of about 50%, the effect on the treated would have been about twice as large.

Appendix Figure 2 plots the coefficient of interest and the 95% confidence interval using alternative bandwidths (up to +/- 90 days). The estimated magnitude of the effect is larger (around 30 days) for smaller bandwidths and stabilizes at around 20 days for bandwidths larger than +/-25. The estimated effects are statistically significant for bandwidths around +/-50 days or larger.

Panel B of Table 1 shows the results stratified by age of the mother at the time of birth of the reference child (median age is 30). We find significant effects for both younger and older mothers, although they tend to be larger and more precisely estimated for the older group.[7] According to our last two specifications (+/- 90 days), both age groups delayed their subsequent birth by more than one month, compared with families of the same age who were not eligible for the paternity leave extension.

We report the results of the difference-in-differences specification (equation (2)) in Apendix Table 4. The first four columns show the results for the full sample, while the last two restrict the sample to mothers older than 30. We present two specifications: one that includes births within eleven weeks before and after March 24 of 2006, 2007 and 2008, and one that includes all births in 2006, 2007 and 2008. The results still show significant effects of the introduction of paternity leave on birth spacing, although the magnitude of the coefficient is reduced by about 50%.

We conclude that families who were eligible for the new paternity leave period took longer to have their next child. In this analysis, we restricted the sample to parents who ended up having (at least) one more child, but it is possible that the likelihood of having a subsequent child might have been affected by the policy. We consider this possibility by estimating equation (1), where the dependent variable is now an indicator for whether the mother had another child in the 2, 4 and 6 years following the birthdate of the reference child. About 7% of the mothers in the relevant window had another child within 2 years, and about 24% did so within the following 4 years (34% within six).

[7] Appendix Table 3 Panel B displays the results for the balance in covariates test corresponding to the sample of women over 30, our preferred bandwidth of +/-77 days, and the same specification employed in Table 1. The results show that the age and education controls are smooth across the threshold.

We report the results of the fertility analysis in Table 2. Panel A includes women of all ages. We find that the introduction of the 13 days of paternity leave significantly reduced the likelihood of having one more child in the following two years, by 7 to 15%, depending on the specification. The effect was still present after 4 years, with eligible households between 1 and 6% less likely to have had another child, although these effects are estimated less precisely. After 6 years, the fertility gap between eligible and ineligible families was still between 1 and 5%.

In Panel B of Table 2 we stratify the results by maternal age. We find no significant effects on subsequent fertility for younger mothers (up to age 30), while the effects are stronger for women who were older than 30 when they had the reference child in 2007. For these older women, additional fertility was reduced by 11-22% in the initial two years, and by a persistent 3-11% (2-9%) after four (six) years. These results are illustrated in Figure 3, where we show the fraction of mothers having another child within 2 years of the reference one, aggregated by week of birth.

The effects for older mothers are statistically significant in the specifications using the broader windows around the threshold. This is illustrated in Appendix Figure 3, where we plot the estimated coefficients and their 95% confidence interval using all possible bandwidths (up to +/- 90 days), for subsequent birth rates after 2 and 4 years. The estimated effects are significantly different from zero for bandwidths over 60 days.

Our results indicate that the reform led eligible families to postpone having an additional child. This delay may have affected subsequent fertility for older women, who may be closer to the end of their fertile cycle. Next, we investigate alternative mechanisms potentially driving these effects. Using different data sets, we quantify the effects of the reform on the labor market outcomes of both parents, their time allocation decisions, fertility preferences, and marital dissolution.

5.2. Labor market outcomes

The introduction of paternity leave could have had effects on labor market outcomes for eligible fathers and/or their partners. Taking family leave may have affected men's work prospects negatively, if for example firms interpreted it as lack of commitment. Alternatively (or additionally), fathers' greater involvement in childcare during their child's first few weeks of life may have had longer-term effects, so that their likelihood of taking unpaid leave or reducing their work hours later on may have increased. With regard to mothers, fathers' increased involvement in childcare may have encouraged them to go back to work earlier after childbirth, and they may have become less likely to take unpaid leave or reduce working hours.

Previous evidence for other countries, while somewhat mixed, suggests that reserving time of the parental leave to fathers, while substantially increasing their take-up rate, does not significantly affect the employment outcomes of either parent (see Ekberg et al. 2013 for Sweden, Bartel et al. 2018 for the US or Patnaik 2015 for Canada).

We analyze the reform's effect on labor market outcomes by estimating equations (1) and (2) using Social Security data. In Appendix Table 5, we use Labor Force Survey data to show that participation rates of mothers and fathers were balanced across the threshold by 2011, remaining so in 2015.[8] This suggests that the Social Security data do not suffer from significant sample selection issues due to differential non-participation. Appendix Table 6 shows evidence that control and treated mothers and fathers are balanced in covariates across the threshold, in terms of their ages, previous number of children, educational attainment, and labor market status before the birth of the reference child.

[8] In fact, the magnitude of the coefficients is not small in 2011, such that mothers from eligible families are more likely to participate. We re-do all of the labor market analysis using only the 2015 sample of the Social Security data, and the results are robust.

The main labor market results are displayed in Table 3. We estimate separate regressions for men and women, and we vary the bandwidth from 3 to 9 months around the threshold. The results for fathers (first three columns) show that their labor market outcomes were unaffected by the two weeks of paternity leave. Eligible fathers were no more or less likely to be working 6, 12, or 24 months after the birth of their child. Their likelihood of taking unpaid family leave was unaffected, as was the probability of holding a part-time position. Their earnings during the months and years following childbirth were unchanged as well. Thus, we find no evidence that the reform had any effect on labor market outcomes for fathers.

We do find some significant effects for mothers. Women whose partners were eligible for paternity leave were more likely to be employed 6, 12, and even 24 months after childbirth. The employment rate of mothers 6 months after having a child is estimated to have increased by 2.5 to 4 percentage points. Given an average of about 55%, our estimates imply an increase of 4.5 to 7.2%. This result is illustrated graphically in Figure 4, where we show maternal employment rates 6 months after the birth of the reference child. Mothers were also significantly less likely to take unpaid family leave. The magnitude of this effect is similar to the one on employment. We also find an increase in the proportion of mothers working part-time.

Put together, our results suggest that some women who would have taken unpaid leave ended up working part-time instead, as a result of the introduction of paternity leave.[9] This is also reflected in higher earnings (a 3 to 5% increase in the first year post-childbirth, given average pre-reform annual earnings of about 10,000 euros).

Since we are using observations as far away from the threshold as 9 months, we address seasonality concerns by reporting difference-in-differences estimates (as described

[9] The effects on employment, unpaid leave and part-time work for mothers are driven by those who were older than 30 in 2007 (results by age not shown).

in equation (2)) in Appendix Table 7. Our results become much less precise. We still find that mothers in eligible households have significantly higher employment rates 6 months after childbirth, but the effect becomes smaller and insignificant by 12 months, and zero by 24 months. The magnitude of the employment effect at 6 months is now estimated to be 4.2 to 4.5%, which remains sizeable given that it implies an effect on the treated that is twice as large.

Overall, the evidence suggests that, even though take-up was high, the two weeks of paternity leave did not affect new fathers' labor market outcomes, or their likelihood of taking extended family leave. On the other hand, mothers whose partners were eligible had significantly higher employment rates six months after childbirth. This increased labor force attachment of mothers may have been driven by the reform increasing fathers' participation in childcare. We test for this possibility in the following section.

5.3. Childcare time and desired fertility

The best available data source for studying childcare time is the Spanish Time Use Survey. The first wave after March 2007 was conducted between October 2009 and September 2010, i.e. about three years after the introduction of the two weeks of paternity leave. Given the low number of observations due to the survey nature of the data, we cannot estimate RD specifications. We report the results of the DiD specification described by equation (2) (using month rather than day of birth), including households who had a child between January and June of 2006, 2007 and 2008 (i.e. 3 months before and after the reform). The relevant coefficient is the one for the interaction of April-June births with the 2007 indicator. We estimate regressions for total daily minutes of childcare time, housework, market work, and any other time use ("residual"), separately for fathers and mothers. Since a number of individuals report zero minutes of childcare, we report both linear and Tobit specifications. The results are reported in Table 4.

The results for fathers suggest that those who would have been eligible for paternity leave in 2007 did almost an hour more childcare per day in 2009-10 compared with ineligible fathers, using families who had a child in the surrounding years as controls. This increase appears to come not from reductions in housework or market work, but in "residual time", most likely leisure and/or sleep.

We find no significant effects on the time-use of mothers. This is consistent with our labor market estimates, that showed no effects on employment or working hours for mothers in the DiD specification two years after the birth of the reference child.

Our time-use analysis provides suggestive evidence that, even though the paternity leave policy did not affect fathers' labor market attachment, it may have affected their involvement in childcare persistently.[10]

So far, we found that most fathers who were eligible for paternity leave took it up. Six months after the birth of the child, mothers in eligible families were more likely to be employed, and three years later fathers were spending more time in childcare. These households delayed the decision to have their next child, and some of them ended up having fewer children. This delay in subsequent fertility may have been driven by mothers' stronger attachment to the labor market, which could have increased the opportunity cost of having another child.

Another potential channel could have been that fathers' increased involvement in childcare lowered their desired fertility. Recent work has shown, in the context of developing countries, that providing information to fathers regarding the cost of having children can lower their desired fertility, as well as families' actual completed fertility, in

[10] Fernández-Cornejo et al. (2016, 2018) also provide suggestive evidence that paternity leave increased fathers' childcare time in Spain, using survey data from Madrid. This result is also in line with the evidence for Germany in Tamm (2018), while Ekberg et al. (2013) do not find effects on fathers' childcare involvement in Sweden.

settings where men desire more children than women and may also enjoy higher bargaining power (Ashraf et al. 2017).

In order to shed light on the plausibility of this channel, we draw from Eurobarometer survey data for 2001, 2006 and 2011, which include questions on desired fertility to a random sample of adults in EU countries. In 2006, men (ages 21 to 40) in Spain reported a higher desired number of children than women, on average. This was not the case in Northern European countries (Sweden, Denmark and Finland, see Figure 8). However, this pattern had reversed by 2011, as displayed in Figure 5. Between 2006 and 2011, men's desired fertility fell, while it increased for women.

In order to test for the statistical significance of this pattern, we estimate regressions for the sample of adult men and women in Spain, including survey data for 2001, 2006, and 2011. The dependent variable is the desired number of children. We include year dummies, and an indicator for male respondents. The explanatory variable of interest is the interaction between the 2011 dummy and the male indicator. The first column of Table 5 only controls for a third-order polynomial in age. Then we add controls for educational attainment and employment status in column 2, and a married dummy plus number of children dummies in column 3. We also estimate a Poisson specification, shown in the final column.

The results of this difference-in-difference analysis show that men reported significantly lower desired fertility in 2011, relative to women. One possible interpretation of this finding, if certainly not the only one, is that men's greater participation in childcare, driven by the introduction of paternity leave in 2007, may have led them to update their fertility preferences downwards. It may be that the extra time with their child made them aware of the full costs of childrearing, and/or that it shifted their preferences from child quantity to quality. This observed shift in preferences may be one reason contributing to the documented decline in subsequent fertility among families who were eligible for the paternity leave permit.

5.4. Marriage stability

A recent study (Avdic and Karimi 2018) finds that an increase in the share of fathers' leave in Sweden led to an increase in divorce in the three years following the birth of the child. If this was also the case in Spain, then parental separation may be an additional (or alternative) channel driving the observed decline in fertility.

We evaluate this possibility by estimating our RD specification (equation (1)) using parental divorce or separation as the outcome variable (with Labor Force Survey data). We estimate the effects on parental separation using cross-sectional data for 2008 to 2013, i.e. one to six years after the birth of the reference child in 2007. The results are reported in Appendix Table 8. The sample includes all women surveyed in a given year who coreside with a child born in the months surrounding March 2007. The first three columns present the results for a binary dependent variable indicating that the woman was divorced, while the last three use a binary variable that takes value 1 if the woman was not cohabiting with a partner at the time of the survey.

We find that women whose partners would have been eligible for paternity leave in 2007 are no more likely to be separated five years after the reform. The evidence suggests a lower divorce propensity three years after childbirth (in 2010), but the effect falls back to zero in 2011, and remains small and insignificant thereafter. Thus, we find no evidence that eligible parents were more likely to break up in the years following the birth of their child, which effectively allows us to rule out this channel.[11]

[11] In Avdic and Karimi (2018) the negative effect of paternity leave on marital stability responds at least in part to a decrease in family income due to mothers' higher use of unpaid leave, which may have increased conflict within the family. This mechanism is

6. Conclusions

We provide evidence suggesting that the introduction of two weeks of paid paternity leave in Spain may have had negative effects on (subsequent) fertility. We show that parents who were (just) entitled to the new paternity leave took longer to have their next child compared to (just) ineligible parents. This delay could explain the observed reduction in subsequent fertility among older women.

We report evidence in support of two (potentially complementary) channels to explain the negative fertility effects. First, fathers' increased involvement in childcare seems to have improved mothers' labor force attachment, as reflected in their higher employment rates after childbirth. This may have increased the opportunity cost of an additional child. We also find that men reported lower desired fertility after the reform. This may have resulted from the paternity leave period raising their awareness about the cost of having children, or from their time with the child increasing their willingness to invest in child quality (versus quantity).

Previous papers found no effects on fertility of increases in paternity leave in Norway (Kotsadam and Finseraas 2001, Dahl et al. 2014, Cools et al. 2015). The Southern European setting is however very different from the Nordic one. Before the 2007 reform, fathers in Spain were making essentially no use of parental leave. Women bore most of the burden of childcare and housework, and their employment rates were relatively low. In this context, a policy that incentivizes fathers' participation in childrearing at the extensive margin may be more effective in increasing women's attachment to the labor force. However, the higher involvement of fathers in childcare activities may have decreased their desired fertility. It remains to be seen whether fertility in Southern European

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unlikely to operate in the same direction in the Spanish case, as we find positive effects on mothers' earnings and a decrease in unpaid leave participation.

countries may fall further or rebound after more extensive reforms of the parental leave system, which may trigger larger changes in social norms and the within-household distribution of market and household work.

More evidence is needed before we can conclude confidently that introducing or extending paternity leave can lead to fewer births, at least in certain contexts, and we should probably be cautious before alerting policy makers to this possibility. Moreover, our analysis focuses only on couples having a child right around the introduction of the policy in Spain, and how their eligibility status affected their subsequent fertility. However, the new paternity leave may also have affected the fertility of all other couples, who would all be eligible in the event of having a child in the future. Those effects, which our identification strategy does not allow us to capture, may have gone in the same or the opposite direction, possibly leading to zero or even a positive total effect on fertility at the national level. This is an interesting avenue for future research.

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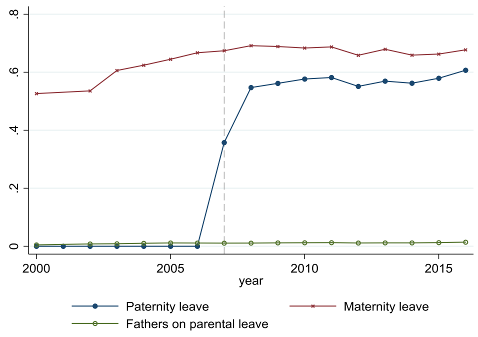
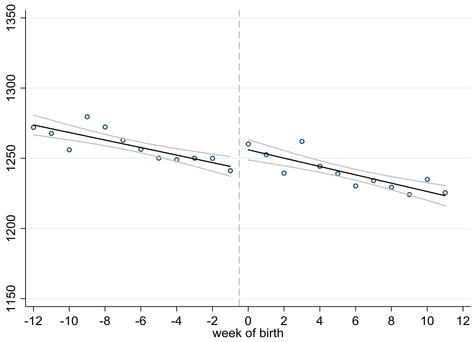
Tables and figures

Figure 1. Number of mothers and fathers taking maternity/paternity leave as a fraction of the annual number of births



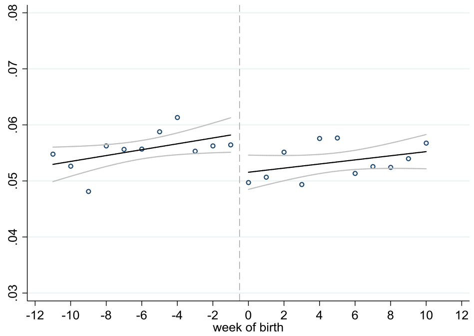
Data sources: Social Security (annual number of leave-takers) and National Statistical Institute (annual number of births).

Figure 2. Days to next birth (birth spacing) by week of birth



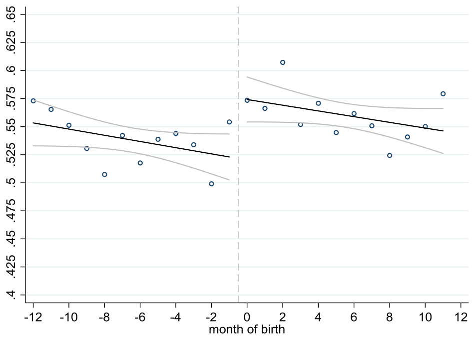
Data source: National Statistical Institute, birth-certificate microdata. The sample includes all women who had a child around the threshold date and who had a subsequent child by the end of 2013. Week of birth is normalized to 0 for women who had the reference child in the week of March 24-30, 2007. The circles are averages within each bin, the black lines are linear fits, and the grey lines are 90% confidence intervals.

Figure 3. Fraction of mothers having another child within the next 2 years (age over 30)



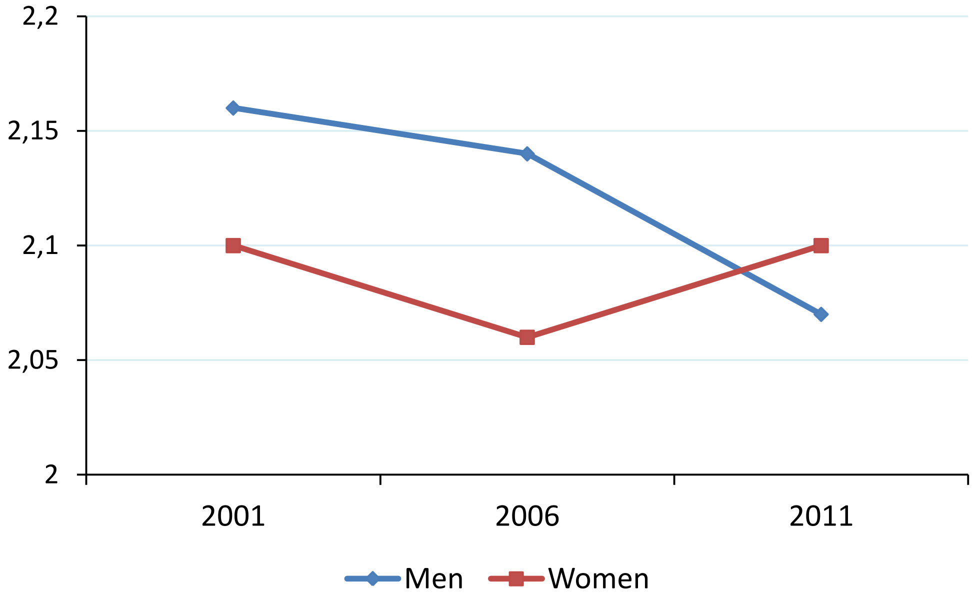
Data source: National Statistical Institute, birth-certificate microdata 2007. Week of birth is normalized to 0 for women having the reference child in the week of March 24-30, 2007. The circles are averages within each bin, the black lines are linear fits, and the grey lines are 90% confidence intervals.

Figure 4. Maternal employment rate 6 months after birth



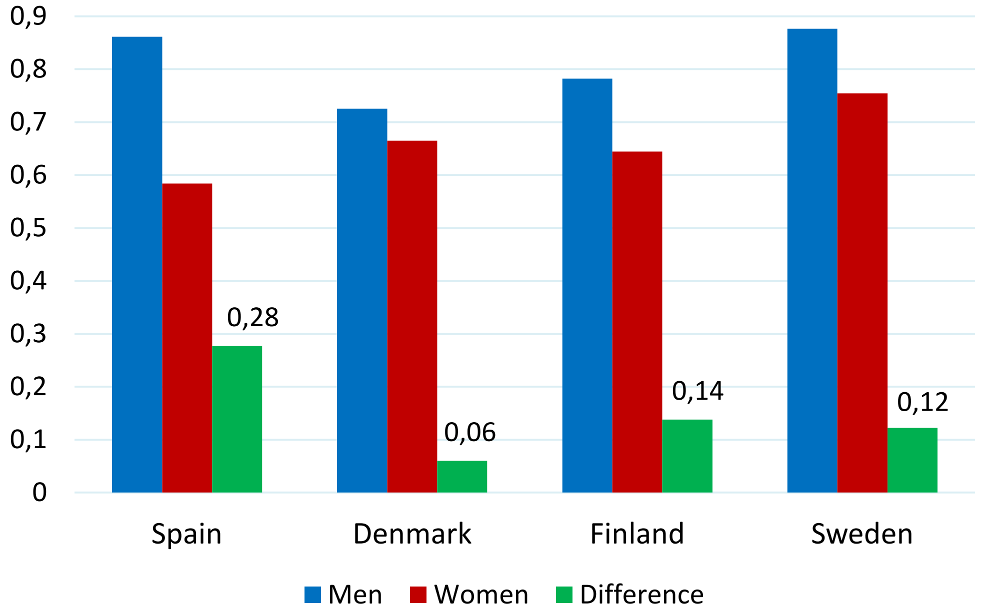
Data source: Social Security Data (*Muestra Continua de Vidas Laborales*). Month of birth is normalized to 0 for children born in April 2007. The circles are averages within each bin, the black lines are linear fits, and the grey lines are 90% confidence intervals.

Figure 5. Desired fertility in Spain, 2001-2011



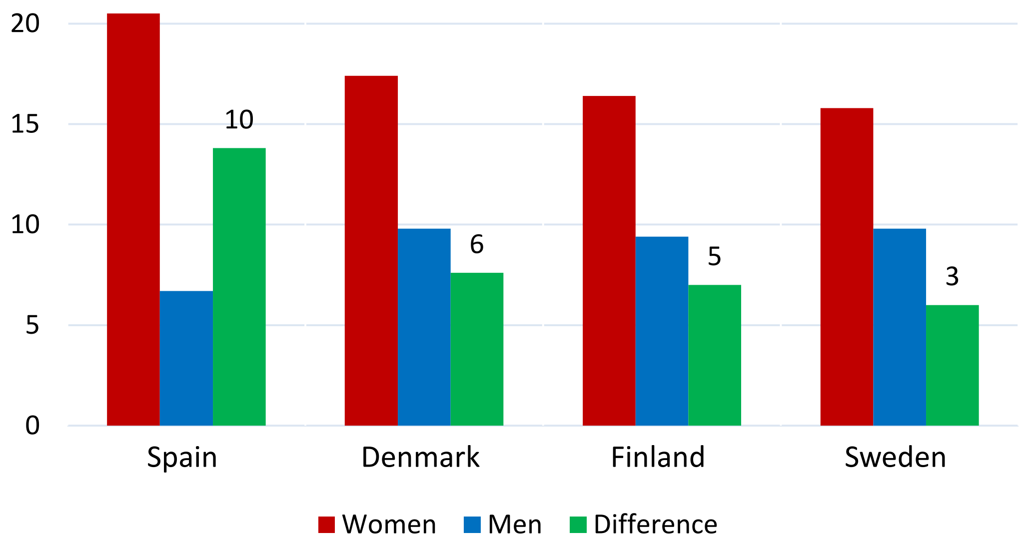
Data source: Eurobarometer. The sample includes all men and women ages 20-40. We report average answers to the question: "For you personally, what would be the ideal number of children you would like to have or would have liked to have had?"

Figure 6. Employment rates in 2006



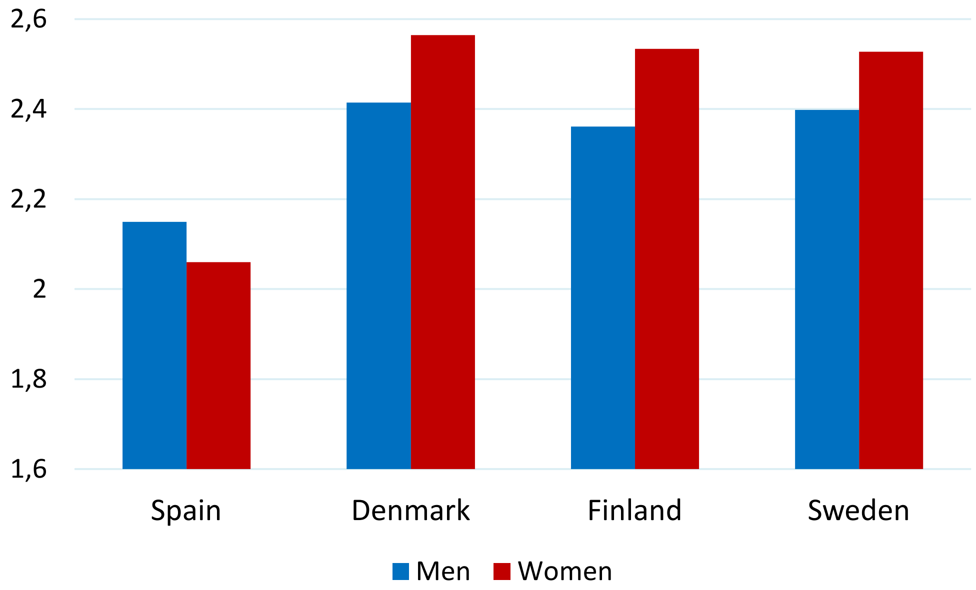
Data source: Eurobarometer survey 2006 (ages 21-40).

Figure 7. Childcare and housework time (%)



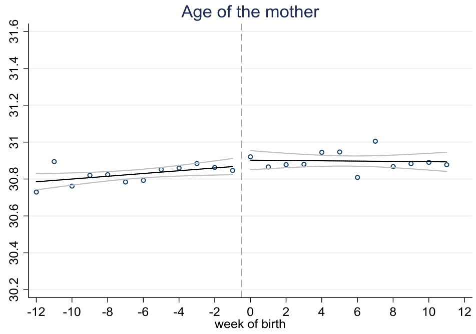
Data source: National Time Use surveys (1999-2003). The bars correspond to the percentage of daily time spent on childcare and housework.

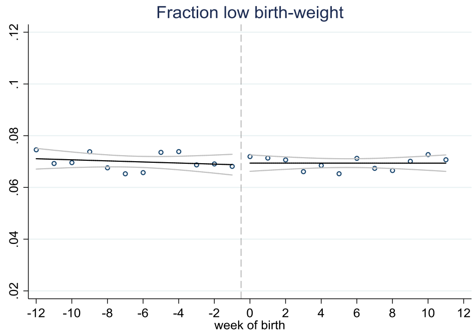
Figure 8. Desired number of children in 2006



Source: Eurobarometer survey, 2006 (ages 21-40). The bars report the averages by gender to the question “For you personally, what would be the ideal number of children you would like to have or would have liked to have had?”.

Figure 9. Balance in covariates at birth





Data source: National Statistical Institute, birth-certificate microdata 2007. Week of birth is 0 for children born in the week of March 24-30, 2007. The circles are averages within each bin, the black lines are linear fits, and the grey lines are 95% confidence intervals.

Table 1.Effect of paternity leave on birth spacing (days to subsequent birth)

Panel A. All ages

Window	Opt. band.	+/- 7 days	+/- 21 days	+/- 56 days	+/- 77 days	+/- 90 days	+/- 90 d. (donut)
Paternity	27.1*	20.7	27.8*	20.5**	16.0**	32.5***	38.2**
Leave	(13.6)	(11.9)	(13.8)	(8.8)	(7.2)	(10.4)	(15.3)
Mean	1,248	1,251	1,249	1,250	1,249	1,249	1,249
Effect as % of mean	2.2%	1.7%	2.2%	1.6%	1.3%	2.6%	3.1%
Linear trends?	Y	N	Y	Y	Y	Y	Y
Quadratic trands?	N	N	N	N	N	Y	Y
Day of the week	Y	N	Y	Y	Y	Y	Y
N	18,329	6,473	19,246	51,628	71,341	83,480	77,007

Panel B. By age

Window	Opt. band.	+/- 7 days	+/- 21 days	+/- 56 days	+/- 77 days	+/- 90 days	+/- 90 d. (donut)
Up to 30							
Paternity	10.2	13.7	17.7	19.4*	11.9	30.9**	41.7**
Leave	(18.3)	(17.7)	(19.5)	(11.6)	(10.5)	(13.7)	(19.6)
N	10,614	3,240	9,675	26,070	36,076	42,562	39,322
Over 30							
Paternity	40.5**	26.4	39.3*	21.5*	20.6**	34.9**	35.7*
Leave	(18.4)	(18.8)	(20.5)	(12.5)	(9.9)	(14.9)	(18.6)
N	10,455	3,233	9,571	25,558	35,265	40,918	37,685
Linear trends?	Y	N	Y	Y	Y	Y	Y
Quadratic trends?	N	N	N	N	N	Y	Y
Day of the week	Y	N	Y	Y	Y	Y	Y

Note: Each coefficient comes from a different regression (standard errors in parentheses). The data come from the birth-certificate microdata files 2006-2013 (National Statistical Institute). The sample includes all mothers who had a child in a certain window (given in column headers) of days around March 24, 2007, and who had another child by the end of 2013. The dependent variable is the number of days between the reference birth and the subsequent one to the same mother; the main independent variable is an indicator for the reference child being born on or after March 24, 2007. Controls include 10 dummies for educational attainment of the mother, and a third-order polynomial in age. The linear (and quadratic) trend in date of birth is interacted with the post-March 24 births indicator. Standard errors are clustered at the date of birth level. The "donut" specification excludes the 7 days right before and after the threshold. The optimal bandwidth is selected via the MSE-RD method (uniform kernel), and is 19.95 for the full sample, and 22.8 and 22.5 for the younger and older subsamples, respectively.

*** $p<0.01$, ** $p<0.05$, * $p<0.1$.

Table 2. Effect of paternity leave on subsequent fertility

Panel A. All ages

Window	Opt. band.	+/- 7 days	+/- 21 days	+/- 56 days	+/- 77 days	+/- 90 days	+/- 90 d. (donut)
			Two years				
Paternity	-0.0097*	-0.0068	-0.0097*	-0.0050	-0.0050**	-0.0084**	-0.0086*
Leave	(0.0053)	(0.0048)	(0.0053)	(0.0031)	(0.0025)	(0.0036)	(0.0049)
Average	0.066	0.064	0.066	0.067	0.067	0.067	0.067
Coeff./average	-14.8%	-10.6%	-14.8%	-7.5%	-7.2%	-12.6%	-12.9%
			Four years				
Paternity	-0.0154**	-0.0066	-0.0120*	-0.0018	-0.0059	-0.0101*	-0.0118
Leave	(0.0068)	(0.0057)	(0.0061)	(0.0044)	(0.0038)	(0.0054)	(0.0085)
Average	0.239	0.238	0.239	0.239	0.239	0.240	0.240
Coeff./average	-6.4%	-2.8%	-5.0%	-0.8%	-2.5%	-4.2%	-4.9%
			Six years				
Paternity	-0.0179*	-0.0093	-0.0129	-0.0023	-0.0071*	-0.0116*	-0.0115
Leave	(0.0091)	(0.0070)	(0.0085)	(0.0050)	(0.0043)	(0.0059)	(0.0108)
Average	0.339	0.336	0.338	0.338	0.337	0.339	0.339
Coeff./average	-5.3%	-2.8%	-3.8%	-0.7%	-2.1%	-3.4%	-3.4%
Linear trends?	Y	N	Y	Y	Y	Y	Y
Quadratic trends?	N	N	N	N	N	Y	Y
Day of the week?	Y	N	Y	Y	Y	Y	Y
N	.	18,174	53,693	144,055	199,558	232,484	214,310

Panel B. By age

Window	Opt. band.	+/- 7 days	+/- 21 days	+/- 56 days	+/- 77 days	+/- 90 days	+/- 90 d. (donut)
				Two years			
Up to 30	-0.0034	-0.0071	-0.0110	-0.0044	-0.0029	-0.0069	-0.0051*
	(0.0060)	(0.0065)	(0.0074)	(0.0043)	(0.0036)	(0.0051)	(0.0071)
Effect as % of mean	-5.1%	-9.0%	-13.6%	-5.4%	-3.5%	-8.4%	-6.2%
Over 30	-0.0084	-0.0063	-0.0084	-0.0057*	-0.0066**	-0.0098**	-0.0120**
	(0.0055)	(0.0053)	(0.0055)	(0.0034)	(0.0029)	(0.0042)	(0.0059)
Effect as % of mean	-12.7%	-12.2%	-16.0%	-10.6%	-12.5%	-18.3%	-22.3%
				Four years			
Up to 30	-0.0151	-0.0066	-0.0121	0.0038	0.0010	-0.0061	-0.0025
	(0.0129)	(0.0111)	(0.0122)	(0.0073)	(0.0063)	(0.0090)	(0.0127)
Effect as % of mean	-6.3%	-2.2%	-4.1%	1.3%	0.3%	-2.0%	-0.8%
Over 30	-0.0093	-0.0058	-0.0114	-0.0067	-0.0108**	-0.0137**	-0.0207**
	(0.0081)	(0.0072)	(0.0085)	(0.0055)	(0.0045)	(0.0065)	(0.0093)
Effect as % of mean	-3.9%	-3.1%	-6.0%	-3.5%	-5.7%	-7.2%	-10.9%
				Six years			
Up to 30	-0.0186	-0.0130	-0.0161	0.0024	-0.0010	-0.0093	-0.0014
	(0.0174)	(0.0167)	(0.0174)	(0.0101)	(0.0083)	(0.0124)	(0.0162)
Effect as % of mean	-5.5%	-3.0%	-3.7%	0.5%	-0.2%	-2.1%	-0.3%
Over 30	-0.0080	-0.0049	-0.0091	-0.0066	-0.0106**	-0.0138**	-0.0220*
	(0.0079)	(0.0071)	(0.0087)	(0.0060)	(0.0051)	(0.0069)	(0.0113)
Effect as % of mean	-2.4%	-1.9%	-3.6%	-2.6%	-4.2%	-5.5%	-8.7%
Linear trends?	Y	N	Y	Y	Y	Y	Y
Quadratic trends?	N	N	N	N	N	Y	Y
Day of the week f.e.	Y	N	Y	Y	Y	Y	Y

Note: Each coefficient comes from a different regression (standard errors clustered by date in parentheses). The data come from birth-certificate microdata files (National Statistical Institute). The sample includes all mothers who had a child in a certain window (given in column headers) around March 24, 2007. The dependent variable is an indicator for the mother having another child within the following 2, 4, or 6 years; the main independent variable is an indicator for the reference child being born on or after March 24, 2007. The linear (and quadratic) trend in date of birth is interacted with the post-March 24 births indicator. Optimal bandwidth is selected using the MSE-SUM method (uniform kernel). The "donut" specification excludes the 7 days right before and after the threshold.
(*** p<0.01. **p<0.05. *p<0.1).

Table 3. Effect of paternity leave on labor market outcomes

	Father			Mother		
	+/-3 months	+/-6 months	+/-9 months	+/-3 months	+/-6 months	+/-9 months
Working after 6 months	-0.003 (0.006)	0.001 (0.009)	0.010 (0.012)	0.040*** (0.009)	0.038*** (0.013)	0.025 (0.017)
Working after 12 m.	-0.003 (0.007)	0.008 (0.010)	-0.001 (0.012)	0.025*** (0.009)	0.031** (0.013)	0.028* (0.017)
Working after 24 m.	-0.008 (0.008)	0.008 (0.012)	0.008 (0.015)	0.011 (0.009)	0.027** (0.014)	0.036** (0.017)
On leave after 6 m.	0.000 (0.001)	0.000 (0.001)	0.000 (0.001)	-0.019*** (0.005)	-0.024*** (0.007)	-0.018* (0.010)
On leave after 12 m.	0.001 (0.001)	0.001 (0.001)	0.001 (0.002)	-0.021*** (0.005)	-0.025*** (0.008)	-0.020** (0.010)
On leave after 24 m.	0.002 (0.001)	0.001 (0.002)	0.002 (0.002)	-0.018*** (0.005)	-0.022*** (0.008)	-0.018* (0.010)
Part-time emp. after 6m.	-0.002 (0.005)	0.000 (0.008)	0.002 (0.010)	0.015** (0.007)	0.024** (0.011)	0.016 (0.014)
Part-time emp. after 12m.	-0.002 (0.006)	-0.001 (0.008)	-0.009 (0.010)	0.010 (0.008)	0.020* (0.012)	0.021 (0.015)
Part-time emp. after 24m.	-0.009 (0.005)	-0.008 (0.008)	-0.013 (0.010)	-0.002 (0.008)	0.011 (0.012)	0.012 (0.015)
Earnings after 6 m.	-42.09 (84.85)	-89.48 (121.71)	-20.64 (154.67)	219.98*** (68.34)	112.80 (99.56)	30.50 (127.46)
Earnings after 12 m.	23.64 (169.06)	-105.10 (242.69)	-78.77 (308.67)	551.33*** (138.18)	381.37* (200.52)	250.98 (256.68)
Earnings after 24 m.	-138.54 (184.35)	-10.07 (6.17)	40.38 (336.11)	380.17** (155.91)	389.89* (226.22)	495.92* (289.84)
N	7,591	15,626	23,477	7,665	15,899	23,853
Liner trend in m	N	Y	Y	N	Y	Y
Quadratic trend	N	N	Y	N	N	Y

Note: Each coefficient comes from a different regression (standard errors in parentheses). The data come from Social Security (*Muestra Continua de Vidas Laborales, 2011-15*). The sample includes all mothers/fathers (aged 16-45 at the time of childbirth) with a child born in a certain window (given in column headers) around March-April 2007. The dependent variable is in the row header; the main independent variable is an indicator for the reference child being born on or after April 2007. The linear (and quadratic) trend in date of birth is interacted with the post-March births indicator. We control for a third-order polynomial in age, occupation, number of children in the house, and indicators for employed, fixed term contract, and public sector employee 3 months before birth.

*** p<0.01. **p<0.05. *p<0.1.

Table 4. Effect of paternity leave on the time-use of fathers and mothers (DiD, +/- 3 months)

	Fathers				Mothers			
Min. per day on:	OLS	Tobit	OLS	Tobit	OLS	Tobit	OLS	Tobit
Child care	45.529	58.320*	52.606	71.948**	-13.080	-10.031	-15.146	-13.190
	(28.283)	(31.468)	(32.769)	(36.168)	(33.060)	(32.648)	(40.640)	(38.675)
Housework	-6.229	3.392	10.005	23.773	42.430	44.706	42.636	45.707
	(28.838)	(34.730)	(34.470)	(40.413)	(27.528)	(27.368)	(37.361)	(36.517)
Work	29.290	39.975	-35.763	-104.885	25.624	54.328	13.452	-56.846
	(70.189)	(118.990)	(78.922)	(134.664)	(50.139)	(117.281)	(68.153)	(158.584)
Residual	-69.095	-69.095	-27.825	-27.825	-54.940	-54.940	-45.489	-45.489
	(59.404)	(56.345)	(69.893)	(65.419)	(40.200)	(38.286)	(54.518)	(50.815)
N	290	290	235	235	313	313	222	222
Age range	16-55	16-55	30-55	30-55	16-55	16-55	30-55	30-55

Note: Each coefficient comes from a different regression (standard errors in parentheses). The data come from the Spanish 2009-10 time-use survey.The sample includes mothers/fathers with a child born in January-June of 2006-08, aged 16-55 on the year of the relevant childbirth.The dependent variable is the number of daily minutes dedicated to each activity; the main independent variable is the interaction between April-June and the indicator for 2007 births. Controls include year dummies, a third-order polynomial in age, an immigrant indicator, number of children younger than 6 in the household, an indicator for the interview being conducted in a work day, education dummies, number of members in the household, and 17 region fixed effects.

*** $p<0.01$. **$p<0.05$. *$p<0.1$.

Table 5. Changes in desired fertility, men vs. women

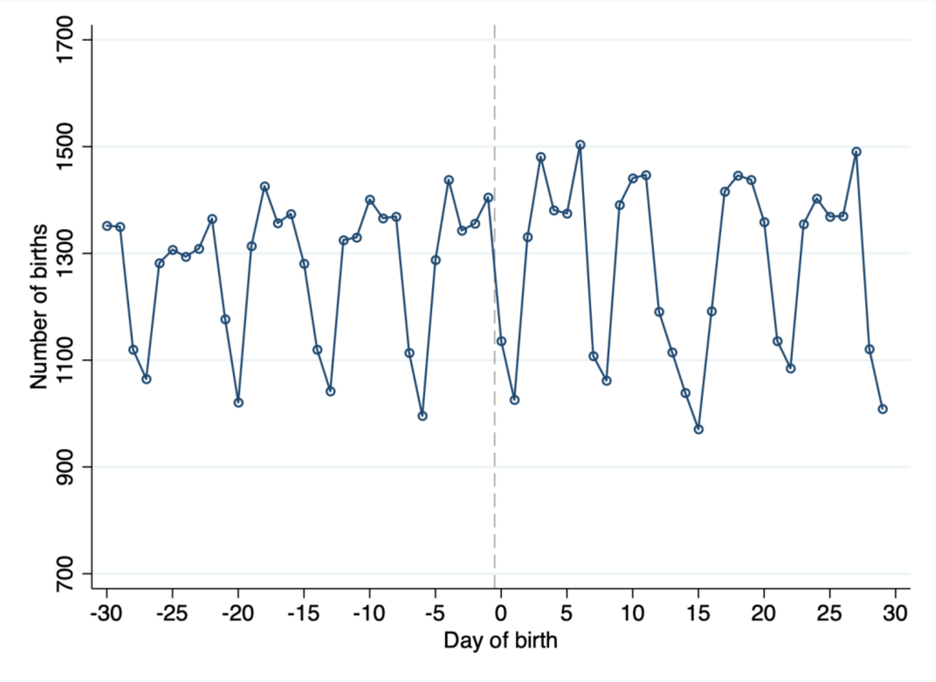
	Basic spec.	LM controls	Full spec.	Poisson
Men*2011	-0.1995 *	-0.2203 *	-0.2214 **	-0.1086 **
	(0.1134)	(0.1146)	(0.1131)	(0.0551)
Men	0.0552	0.0546	0.1955 *	0.0988 *
	(0.0762)	(0.0777)	(0.1089)	(0.0556)
N	871	871	868	868
Age polynomial	Y	Y	Y	Y
Educ. & emp.	N	Y	Y	Y
Married & n. children controls	N	N	Y	Y

Note: Each coefficient comes from a different regression (standard errors in parentheses). The data come from the Spanish 2001, 2006 and 2011 Eurobarometer survey. The sample includes men and women aged 23 to 40.The dependent variable is the desired number of children; the main independent variable is the interaction between an indicator for men and the 2011 dummy. All specifications include year dummies and a third-order polynomial in age. Full controls include a married dummy, three education indicators, an employment indicator, and dummies for number of children interacted with male.

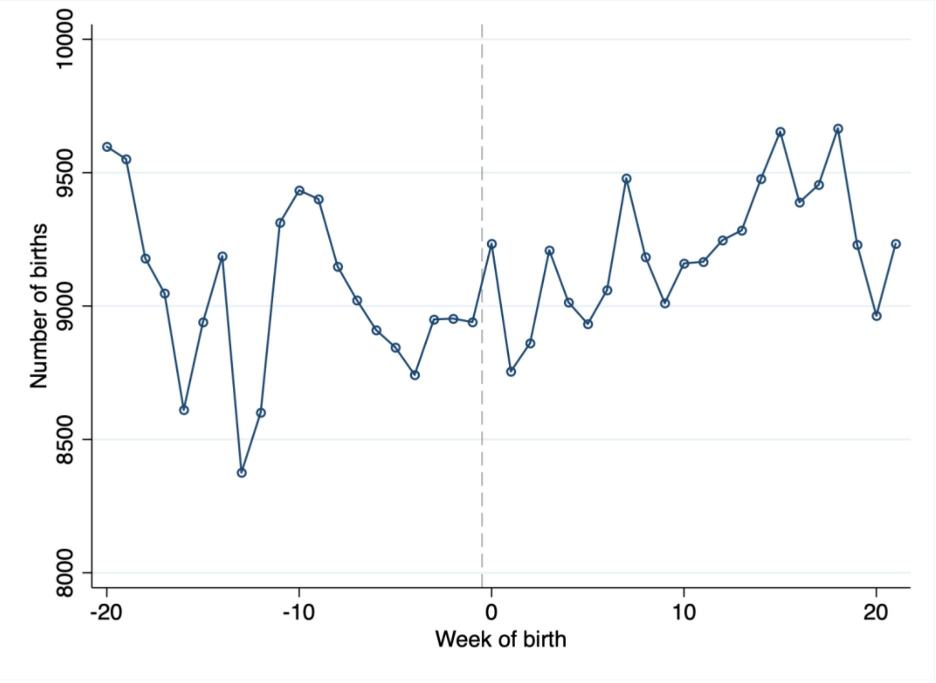
*** p<0.01. **p<0.05. *p<0.1.

Appendix Figure 1. Daily and weekly number of births around the introduction of paternity leave

Panel A. Daily number of births

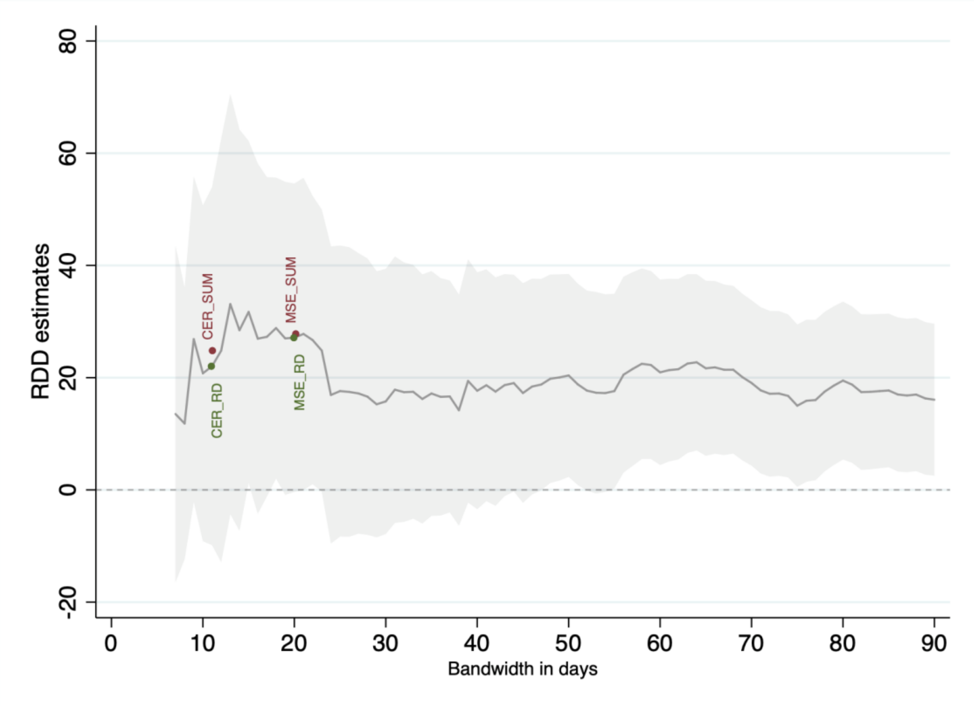


Panel B. Weekly number of births



Source: Birth-certificate data, 2006-2007.
Note: Day (week) of birth is normalized to 0 for March 24 (the week of March 24 to March 30), 2007.

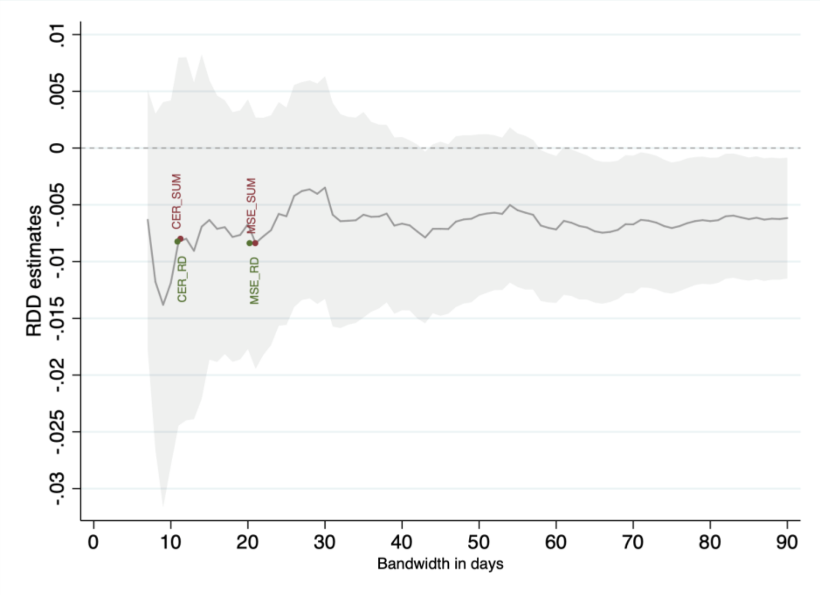
Appendix Figure 2. Effect of paternity leave on birth spacing (days to subsequent birth), different bandwidths



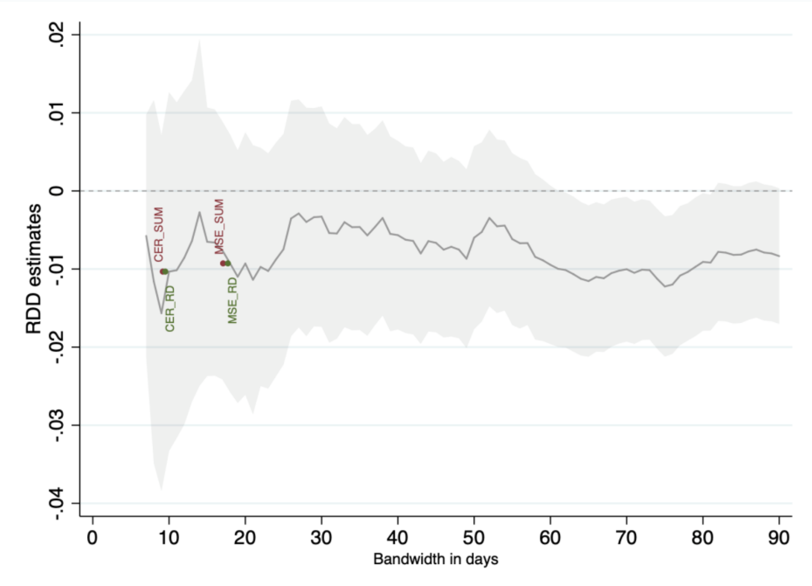
Note: The line displays the results of regressions varying the bandwidth (from 7 to 90 days before and after the threshold), while the shaded area shows the 95% confidence interval for each coefficient. The data come from the birth-certificate microdata files 2006-2013 (National Statistical Institute). The sample includes all mothers who had a child in a certain window (given in horizontal axis) of days around March 24, 2007, and who had another child by the end of 2013. The dependent variable is the number of days between the reference birth and the subsequent one to the same mother; the main independent variable is an indicator for the reference child being born on or after March 24. 2007. Controls include 10 dummies for educational attainment of the mother, and a third-order polynomial in age. A linear trend in date of birth is interacted with the post-March 24 births indicator. Standard errors are clustered at the date of birth level. We show the optimal bandwidth for four selection methods (MSE-RD, MSE-SUM, CER-RD, and CER-SUM). The MSE- methods specify one common mean squared error (MSE)-optimal bandwidth selector. The CER- methods specify one common coverage error-rate (CER)-optimal bandwidth selector. The -RD methods specify the optimal bandwidth selector for the RD treatment-effect estimator, while the -SUM methods specify it for the sum of regression estimates.

Appendix Figure 3. Effect of paternity leave on subsequent fertility, mothers 30 or older, different bandwidths

Panel A. Effect on fraction of mother having another child within 2 years



Panel B. Effect on fraction of mothers having another child within 4 years



Note: The lines display the results of regressions varying the bandwidth (from 7 to 90 days around the threshold). The shaded areas show 95% confidence intervals. The data come from birth-certificate microdata files (National Statistical Institute).The sample includes all mothers aged 30 or older who had a child in a certain window (in horizontal axis) around March 24, 2007.The dependent variable is an indicator for the mother having another child within the following 2 or 4 years; the main independent variable is an indicator for the reference child being born on or after March 24, 2007. A linear trend in date of birth is interacted with the post-March 24 indicator. We show the optimal bandwidth for four selection methods (MSE-RD, MSE-SUM, CER-RD, and CER-SUM). The MSE- methods specify one common mean squared error (MSE)-optimal bandwidth selector. The CER- methods specify one common coverage error-rate (CER)-optimal bandwidth selector. The -RD methods specify the optimal bandwidth selector for the RD treatment-effect estimator, while the -SUM methods specify it for the sum of regression estimates.

Appendix Table 1.Parental leave reforms in Spain

	March 1980 Statute of Rights for Workers ***March 1984*** Law 30/1984 for the reform of the Public Service	***March 1989*** Law 3/1989 to extend maternity leave to 16 weeks and to promote gender equality at the work place	***November 1999*** Law 39/1999 to promote work and family life	***March 2007*** 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 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.

Appendix Table 2. Bunching in number of births at the threshold

Window	+/- 7 days	+/- 14 days	+/- 21 days	+/- 42 days	+/- 56 days	+/- 77 days	+/- 90 days
Daily n. of births	42	72	40	12	9	32	20
	(91)	(54)	(39)	(31)	(26)	(21)	(21)
Log n. of births	0.0316	0.0554	0.0314	0.0062	0.0200	0.0214	0.0202
	(0.074)	(0.0414)	(0.0290)	(0.0235)	(0.047)	(0.0165)	(0.0238)
Linear trend	N	Y	Y	Y	Y	Y	Y
Quadratic trend	N	N	N	N	N	N	Y
Day of the week f.e.	N	Y	Y	Y	Y	Y	Y
N.obs.	14	28	42	84	112	154	172

Source: Birth-certificate data, 2007.
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.

*** $p<0.01$. **$p<0.05$. *$p<0.1$.

Appendix Table 3. Balance in covariates around the paternity leave extension

Panel A. Birth outcomes and parents' characteristics, alternative specifications

Window	+/- 7 days	+/- 21 days	+/- 56 days	+/- 77 days	+/- 90 days
Weeks of gestation	-0.0006	0.0084	0.0199	0.0161	-0.0033
	(0.0290)	(0.0306)	(0.0193)	(0.0167)	(0.0237)
Prematurity	0.0035	0.0015	-0.0006	-0.0005	0.0009
	(0.0033)	(0.0040)	(0.0026)	(0.0022)	(0.0032)
Birth weight	-0.0034	-0.0055	0.0006	0.0018	-0.0018
(in logs)	(0.0033)	(0.0034)	(0.0020)	(0.0017)	(0.0024)
Low birth-weight	0.0038	0.0038	0.0004	-0.0002	0.0033
	(0.0041)	(0.0040)	(0.0025)	(0.0022)	(0.0032)
Mortality 24h.	-0.0106	-0.1735	-0.0293	-0.0912	0.0159
	(0.2618)	(0.2987)	(0.2013)	(0.1725)	(0.2523)
Female	0.0035	0.0028	0.0066	0.0003	0.0054
	(0.0075)	(0.0072)	(0.0045)	(0.0040)	(0.0057)
Mother's age	0.0667	0.0796	-0.0107	0.0338	0.0088
	(0.1519)	(0.1110)	(0.0608)	(0.0519)	(0.0765)
Mother college	0.0016	-0.0001	0.0043	0.0001	0.0162 **
educated	(0.0101)	(0.0070)	(0.0049)	(0.0046)	(0.0070)
Mother out of	-0.0009	-0.0023	-0.0042	-0.0083 **	-0.0027
the labor force	(0.0061)	(0.0050)	(0.0035)	(0.0033)	(0.0045)
Mother foreign-born	-0.0002	0.0006	-0.0091 *	-0.0184 ***	-0.0023
	(0.0101)	(0.0079)	(0.0049)	(0.0045)	(0.0057)
Father's age	0.0432	0.0681	-0.0046	0.0044	0.0230
	(0.0875)	(0.0677)	(0.0471)	(0.0399)	(0.0563)
Father college	-0.002	-0.0044	0.0012	-0.0032	0.0087
educated	(0.0086)	(0.0056)	(0.0043)	(0.0040)	(0.0056)
Linear trends?	N	Y	Y	Y	Y
Quadratic trends?	N	N	N	N	Y
Day of the week?	N	Y	Y	Y	Y
N	18,119	53,522	143,623	198,947	222,513

Note: Each coefficient comes from a different regression (standard errors in parentheses). The dependent variable is the row header; the main independent variable is an indicator for families with children born on or after March 24, 2007. The data come from the 2007 birth-certificate microdata file, National Statistical Institute.The sample includes all births that took place in a certain window (given in column headers) of days around March 24, 2007.The linear trend in date of birth (the running variable, centered at 0 in March 24, 2007) is interacted with the main explanatory variable. Standard errors are clustered at the date of birth level.

Panel B. Controls included in main specification

Covariates	+/- 77 days
Mother's age	-0.0344
	(0.0543)
Illiterate	-0.0015
	(0.0018)
Less than 5 years of education	-0.0007
	(0.0022)
Incomplete primary education (> 5 years)	-0.0037
	(0.0051)
Complete primary education	-0.0123
	(0.0078)
High school	0.0096
	(0.0059)
Short vocational education	-0.0003
	(0.0043)
Long vocational education	0.0006
	(0.0066)
Short college degree	0.0029
	(0.0080)
Long college degree	0.0065
	(0.0102)
PhD	-0.0011
	(0.0031)

Note: Sample restricted to women over 30 having a child on the 77 days before and after the reform. Each coefficient comes from a different regression, where the dependent variable is the variable in each row header, and the main independent variable is an indicator for the reference child being born on or after March 24, 2007. The education variables refer to the mother. All regressions control for a linear trend in the date of birth (interacted with the post-March 24 birth indicator) and day of the week fixed-effects. Robust standard errors in parentheses, clustered by date of birth.
*** p<0.01, ** p<0.05, * p<0.1

Appendix Table 4. Effect of paternity leave on birth spacing (days to subsequent birth), difference-in-difference specification

	All ages				Older mothers (>30)	
Window	**+/- 77 days**	**+/- 77 days**	**Full year**	**Full year**	**+/- 77 days**	**Full year**
Paternity leave	7.61 (4.65)	8.02 * (4.70)	6.97 ** (3.51)	6.98 ** (3.47)	15.47 ** (6.25)	12.39 ** (4.86)
Mean	1,251	1,251	1,221	1,221	1,194	1,165
Effect as % of mean	0.6%	0.6%	0.6%	0.6%	1.3%	1.1%
N	204,756	204,756	506,936	506,936	100,857	238,058
Day of the week	N	Y	N	Y	N	Y
Linear trend in d	Y	Y	Y	Y	Y	Y
Cubic trend in d	N	Y	N	Y	N	Y

Note: Each coefficient comes from a different regression (standard errors clustered by date of birth in parentheses). The data come from birth-certificate microdata files (National Statistical Institute) for 2006, 2007 and 2008.The sample includes all mothers who had a child+/- 77 days around March 24 of each of the three years, or at any time during the three years (depending on the column), and who had another child by the end of 2013.The dependent variable is the number of days between the reference birth and the subsequent one to the same mother; the main independent variable is the interaction between the reference child being born on or after March 24, and an indicator for 2007 births. Controls include year dummies, a post-March 24 births indicator, 10 dummies for educational attainment of the mother, and a third-order polynomial in age.

*** p<0.01, **p<0.05, *p<0.1.

Appendix Table 5. Effect of paternity leave on labor market participation 4 and 8 years later

	2015			2011		
	+/-3 months	+/-6 months	+/-9 months	+/-3 months	+/-6 months	+/-9 months
Mothers	-0.008	0.001	-0.000	0.035**	0.026	0.026
	(0.014)	(0.021)	(0.027)	(0.014)	(0.020)	(0.026)
N	3,302	6,540	9,916	3,299	6,755	10,175
Fathers	0.002	0.006	0.004	0.004	0.012	0.020*
	(0.006)	(0.010)	(0.013)	(0.006)	(0.009)	(0.012)
N	2,746	5,462	8,281	2,959	6,062	9,099
Liner trend in m	N	Y	Y	N	Y	Y
Quadratic trend in m	N	N	Y	N	N	Y

Note: Each coefficient comes from a different regression (standard errors in parentheses). The data come from the Labor Force Survey (2011 and 2015).The sample includes all mothers/fathers (aged 15-45 at the time of childbirth) with a child born in a certain window (given in column headers) around March-April 2007.The dependent variable is an indicator for labor force participation; the main independent variable is an indicator for the reference child being born on or after April 2007. The trend in month of birth is always interacted with the post-March births indicator. Control variables include athird order polynomial in age, education dummies, and an immigrant indicator.

*** p<0.01. **p<0.05. *p<0.1.

Appendix Table 6. Balance in covariates, Social Security data

	Age	N. of children	High educ.	Medium educ.	Low educ.	Employed	Permanent contract	Public employee
Mothers								
Paternity leave	-0.329*	-0.009	-0.006	0.026*	-0.021	-0.000	0.006	0.005
	(0.176)	(0.047)	(0.008)	(0.016)	(0.015)	(0.009)	(0.010)	(0.008)
N	15,899	15,899	15,899	15,899	15,899	15,899	15,899	15,899
Fathers								
Paternity leave	-0.128	0.047	-0.009	0.010	-0.001	0.015	0.009	0.000
	(0.176)	(0.047)	(0.008)	(0.016)	(0.016)	(0.010)	(0.012)	(0.007)
N	15,626	15,626	15,626	15,626	15,626	15,626	15,626	15,626

Note: Each coefficient comes from a different regression (standard errors in parentheses). The data come from Social Security (*Muestra Continua de Vidas Laborales, 2011-15*). The sample includes all mothers/fathers (aged 16-45) with a child born in a +/- 6-month window around March-April 2007.The dependent variable is in the column header; measured before the reference birth. The main independent variable is an indicator for the reference child being born on or after April 2007. All specifications include a linear trend in month of birth, interacted with the post-March births indicator.

*** p<0.01. **p<0.05. *p<0.1.

Appendix Table 7. Effect of paternity leave on labor market outcomes, difference-in-difference specification

	Father		Mother	
	+/-3 months	Full year	+/-3 months	Full year
Working after 6 months	0.002	0.001	0.023*	0.025**
	(0.008)	(0.007)	(0.014)	(0.012)
Working after 12 months	-0.002	-0.021***	0.013	0.010
	(0.009)	(0.008)	(0.014)	(0.012)
Working after 24 months	-0.012	-0.015*	-0.003	-0.005
	(0.011)	(0.009)	(0.014)	(0.012)
On unpaid leave after 12 m.	-0.000	0.000	-0.015	-0.004
	(0.001)	(0.001)	(0.010)	(0.008)
On unpaid leave after 24 m.	0.001	0.001	-0.012	-0.003
	(0.002)	(0.001)	(0.010)	(0.009)
Part-time emp. after 6 months	-0.010	-117.748	0.016	0.003
	(0.008)	(121.225)	(0.013)	(0.011)
Part-time emp. after 12 months	-0.014*	-0.004	0.006	0.005
	(0.008)	(0.007)	(0.012)	(0.010)
Earnings after 6 months	-16.709	5.962	-117.748	-6.335
	(127.764)	(109.475)	(121.225)	(101.881)
Earnings after 12 months	10.817	-25.495	-109.297	115.621
	(254.475)	(217.416)	(245.314)	(205.600)
Age range	>30	>30	>30	>30
N	16,014	31,485	13,571	26,639

Note: Each coefficient comes from a different regression (standard errors in parentheses). The data come from Social Security (*Muestra Continua de Vidas Laborales, 2011-15*).The sample includes all mothers/fathers (aged 16-45 at childbirth) with a child born in +/- 3 month window around March-April 2006, 2007 and 2008 (1st and 3rd columns), or with a child born at any point during 2006-08 (2nd and 4th columns).The dependent variables are in the row header; the main independent variable is the interaction between an indicator for the reference child being born on or after April, and a 2007 indicator. We control for year dummies, a third-order polynomial in age, occupation dummies (high, medium, and low), number of children in the house, and indicators for employed, fixed term contract, and public sector employee 3 months before birth.

*** p<0.01. **p<0.05. *p<0.1.

Appendix Table 8. Effect of paternity leave on separation and divorce

	Divorced			Not cohabiting		
	+/-3 months	+/-6 months	+/-9 months	+/-3 months	+/-6 months	+/-9 months
2008	0.009	0.003	0.008	0.013	0.015	0.026
	(0.007)	(0.009)	(0.011)	(0.009)	(0.014)	(0.017)
2009	0.004	0.017*	0.024**	-0.003	0.014	0.011
	(0.007)	(0.009)	(0.012)	(0.010)	(0.014)	(0.017)
2010	-0.008	-0.025**	-0.028**	-0.022**	-0.036***	-0.035**
	(0.006)	(0.010)	(0.013)	(0.009)	(0.013)	(0.017)
2011	-0.003	-0.014	-0.015	-0.010	-0.025*	-0.029*
	(0.007)	(0.010)	(0.013)	(0.010)	(0.014)	(0.017)
2012	-0.000	0.004	0.014	-0.001	-0.002	0.003
	(0.008)	(0.012)	(0.015)	(0.011)	(0.016)	(0.020)
2013	0.005	0.010	0.009	0.007	0.020	0.006
	(0.009)	(0.013)	(0.017)	(0.011)	(0.016)	(0.021)
Liner trend in m	N	Y	Y	N	Y	Y
Quadratic trend in m	N	N	Y	N	N	Y

Note: Each coefficient comes from a different regression (standard errors in parentheses). The data come from the Labor Force Survey (year indicated in row headers). The sample includes all mothers (aged 16-45 at the time of childbirth) with a child born in a certain window (given in column headers) around March-April 2007.The dependent variable is an indicator for being divorced or not cohabiting with a partner at the time of the survey; the main independent variable is an indicator for the reference child being born on or after April 2007. The trend in month of birth is always interacted with the post-March births indicator. Control variables include a third order polynomial in age, education dummies, an immigrant indicator, number of children born before the reference child, and quarter fixed effects.

*** $p<0.01$. **$p<0.05$. *$p<0.1$.