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Men's Equality at Home for Women's Equality at Work: The Spanish Case

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Abstract

Maternity leave has been at the center of family reconciliation policies for many decades. Despite being fundamental for the wellbeing of mother and baby, inequalities for women in the labor market arise. Therefore, in recent years countries have focused on increasing the duration of paternity leave, and even bringing it into line with that of mothers. In Spain, this point was reached in less than 15 years after the extension of paternity leave to two weeks. The changes that this policy generated have been studied in the short term, but the impact on families today is virtually unknown, being the latter the main focus of this project. Based on Spanish data from the 2018 Fertility Survey, using a regression discontinuity design approach, it is found that father's involvement in numerous caregiving tasks increased even 10 years later. As less than the sole responsibility felt on mothers, they were more likely to have a full-time job and had higher satisfaction with childcare sharing. Using the same database, the short-term effects of the expansion of paternity leave to four weeks in 2017 are analyzed. The effects found are much smaller, which may be due to the fact that the individuals had lower gender norms and that father's involvement in childcare was already higher.

Key words: Paternity leave, gender equality, Spanish policies, long-term father involvement

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1 Introduction

United Nations Sustainable Development Goal number 5 is to achieve gender equality and empower all women and girls (United Nations, 2022). Maternity leave can contribute to gender inequality and the gender pay gap. Employers may be less likely to hire or promote women of childbearing age, while mothers who take time off for childcare may face reduced earning potential and career advancement opportunities, leading to a wider gender pay gap over time. An extension of maternity leave tends to increase female participation in the labor market in exchange for lower wages, a lower presence of women in high-skilled occupations and a more traditional division of tasks within the household (Farré, 2016; Schönberg & Ludsteck, 2014). However, maternity leave cannot be abolished due to these negative effects, as the positive ones are essential to improve the well-being of the mother and the baby. Maternity leave has been shown to have positive effects on mother's physical and mental health (Aitken et al., 2015; Avendano et al., 2015; Van Niel et al., 2020), the success of breastfeeding (Guendelman et al., 2009), or taking the baby to the doctor for visits and being up to date in immunizations (Berger et al., 2005; Heymann et al., 2017). Additionally, early return to work from the mother translates into children's lower cognitive test scores years later (Baydar & Brooks-Gunn, 1991; Ruhm, 2004), behavior problems (NICHD, 1998) and higher infant mortality (Nandi et al., 2016; Rossin, 2011).

To overcome the inequalities that maternity leave can give rise to, many countries as South Korea or France (see the full list in The World Bank (2023)), have centered the focus on paternity leaves following Sweden's first steps in 1974. These promote gender equality by encouraging men to take a more active role in child-rearing and reducing the burden of caregiving primarily falling on women. At the European level, it has also been at the center of the political debate. In June 2019, the European Parliament adopted a directive fixing the minimum requirement of a nontransferable paternity-leave quota of 2 weeks (Council Directive 2019/1158). Dearing (2016) constructed an indicator that defines a well-paid ideal leave model of 14 months, with half reserved for fathers. Comparing European paternity leaves to the desired blueprint, she concludes that most align better with the ideal overall duration and pay, rather than the desired share reserved for fathers. This reflects that the path for improvement is the increase in the share of leave earmarked for dads, which makes the uptake of fathers increase significantly (Lappegård, 2012). Nevertheless, the design of parental leave is key for its impact on gender equality, depending for example on duration, income replacement, non-transferability (Castro-García & Pazos-Moran, 2016) and the interrelation of the various leaves. For example, paternity leave will have a greater impact when the father uses his weeks alone than jointly with the mother (Wall & Escobedo, 2013).

A main driver of changes in gender norms are intergenerational effects. Many studies have found that individuals' housework provision is predicted by the parental division of housework during their childhoods, mother's employment, and mother's gender role ideology (Cordero-Coma & Esping-Andersen, 2018; Cunningham, 2001a, 2001b; Moen et al., 1997). Children's attitudes and behaviours have been found to be more gender equalitarian if their fathers were entitled to paternity leaves (Farré et al., 2023; Kotsadam & Finseraas, 2013). Beliefs in gender roles are changing but the effect of intergenerational effects takes time. Policies are needed to speed up the arrival of more gender equality, keeping in mind that they don't only affect present beliefs but also can have positive spillover effects over other generations (Dahl & Gielen, 2021; Farré et al, 2023).

The motivation for this paper arises from the curiosity awakened by the long-term effects that paternity leaves could have on several outcomes after 10 years, being the ones of main interest childcare and housework responsibility, number of children and the amount desired, and the wellbeing of both fathers and mothers. Due to the recent history of these permits, there is a gap in the literature on their long-term effects, being there only 3 papers that focus on it. Kotsadam & Finseraas (2011) use Norwegian data from 14 years after the introduction of a paternity leave and find it had a causal effect on reducing conflicts over housework, equally sharing the task of washing clothes and increasing the support for public provision of childcare. For the same country, Cools et al. (2015) find significant effects on school performance, which improved for 14- to 16-year-olds as a result of the paternity leave introduction. On the contrary, the probability of being married and several labor market outcomes did not vary. Lastly, Korsgren & van Lent (2022) use several paternity leave increases in Europe, having some of them occurred 13 years earlier, to find that they increased life satisfaction for both parents but 30% more for mothers than fathers. Job satisfaction and work-life balanced did not change, which suggests that the increase in general satisfaction comes from other factors.

The previous results imply an increase in long-term child responsibility in line with Becker (1991) and Becker et al. (1985), who state that even a short period of paternity leave might affect the evolution of household roles. A small change in initial comparative advantages can be sufficient to generate a larger impact in the longer run. On the contrary, Farré (2016) states that long-term effects (what she considers 3 years) of the periods of leave reserved to men have not been found in many cases (Ekberg et al., 2013; Kluve & Tamm, 2013; Schober, 2014), as culture and believes evolve slowly over time (Alesina et al., 2013; Farré and Vella, 2013). The main goal of this study is to check if long-term child responsibility arises with the introduction of a paternity leave of 2 weeks in Spain in 2007, as Becker's ideas state. For this to be true, an increase of leave of absence for childcare by fathers who were eligible for the leave is expected. The leave of absence for childcare by fathers consist of an unpaid period without working to take care of a child under 3 years of age, while retaining the right to return to work. A rise in housework and childcare provided by the dads 10 years later is also anticipated because of their greater long-term commitment, and therefore an improvement in the mother's satisfaction regarding the division of these two tasks.

The second objective of this paper is to test whether the effects that Farré & González (2019) find in Spain in the short run hold in the long run for the 2007 reform, as well as whether the short run effects for the 2017 reform are the same they found for the previous one. It can be expected that this last reform might have generated less changes in fathers' behaviors due to the gender convergence that has occurred over the last decades as a result of the evolution of gender preferences and norms (Kleven et al., 2022). Finally, one variable whose relationship with paternity leave has never been studied is the number of children desired. Findings on the effect of paternity leave on fertility are mixed (Cools et al., 2015; Hart et al., 2019; Lappegård & Kornstad, 2020; Raute, 2019), while there is greater agreement that birth spacing increases as a result (Cygan-Rehm, 2016; Farré & González, 2019; Fontenay & Tojerow, 2020). It will be of interest to study whether the desired number of children behaves similarly to other fertility-related outcomes. A difference between both variables can easily arise for two reasons. First, even if families desired to have more kids after the introduction of the reform, the difficult macroeconomic situation that arose one year after due to the economic crisis of 2008 might not have allowed families to materialize their wish to increase their family size. Additionally, if the paternity leave made only one member of the family desire more kids, it might not have translated into having more children due to differences with their partner's preferences.

Some authors (Bertrand, 2011; Kleven et al., 2019; Kleven et al., 2022) defend that gender convergence captures the evolution of gender preferences and norms, and it is not driven by family policies themselves. Despite that, there exists a wide literature that finds beneficial effects of paternity leaves right after their introduction for several outcomes. First, some research focuses on gender wage gap and labor market outcomes of both parents. The decrease in pay gap is usually driven by a positive effect on mothers' labor supply and earnings, while not by a negative effect of fathers' labor outcomes. Examples of this are studies from Canada (Patnaik, 2019), Denmark (Andersen, 2018;

Druedahl et al., 2019), Germany (Frodermann et al., 2023; Tamm, 2019) or Spain (Farré & González, 2019). Nonetheless, results are sometimes mixed and depend on the country and data used. Regarding two of the leaders in the introduction of paternity leaves, Sweden and Norway, some papers find no effect on earnings and work hours of both parents (Cools et al., 2015; Ekberg et al., 2013; Hart et al., 2019). Studying the same reforms, others conclude that father's yearly income was reduced due to spending more time on home production, and that positive effects on mothers' earnings arose (Johansson, 2010; Rege & Solli, 2013).

Secondly, paternity leave increases the provision of housework by men, as found in the European Union (Meil, 2011) and many countries like the US (Petts et al., 2020), Norway (Kotsadam & Finseraas, 2011, Rege & Solli, 2013) or Spain (Farré & González, 2019; González & Zoabi, 2021). This arises through two mechanisms, father-child bonding and prevention of exclusive expertise in childcare from the mother (Tanaka & Waldfogel, 2007). Men's equality in the home is needed for women's equality in the workplace, as time spent on household chores affects working time and vice versa. It must be kept in mind that the connection between both time uses might not be linear or even directly causal, but they are related in some way. The smallest differences between time spent by men and women on caring and household chores are found in OCDE countries with the smallest gender gaps in employment rate (OECD, 2016). Even if the gender gap in unpaid work has been proven to be smaller for younger couples (OECD, 2016), the arrival of parenthood has been found to be one of the biggest explanations of sharing paid and unpaid household and childcare work the traditional way (Baxter et al., 2008; Grunow et al., 2012). That emerges because of changes in cognitive beliefs that make their attitudes more traditional (Baxter et al., 2015) and the institutional arrangements that promote traditional family divisions, being this last component what can become more gender equal through policies.

Finally, few studies have focused on the effect that paternity leave has on the wellbeing of all family members. Regarding the health of the mother, maternal depression intensity has been found to be associated with lack of parental involvement (Séjourné et al., 2012), and paternal access to workplace flexibility reduced the risked of mothers experiencing physical postpartum health complications and improved their mental health (Persson & Rossin-Slater, 2019). Another welfare increasing result are improvements in relationships among family members. Closer relationships between fathers and children arise as one of the improvements from paternity leave, as reported by both parties (Petts et al., 2020; Wilson & Prior, 2010). Paternity leaves also help to strengthen families by nurturing higher quality father-mother relationships as it signals a greater investment in family life, translating into an increase in spousal stability (Norman et al., 2018; Petts et al., 2020). Notwithstanding, in this topic mixed results arise. In Norway union stability was unaffected by an increase of the paternity leave (Hart et al., 2019), while it led to an increase in divorces in Sweden (Avdic & Karimi, 2018) and Spain (Farré & González, 2019). Additionally, there also exist associations between paternity leave and parents' reports of higher relationship satisfaction and less relationship conflict, by the age the babies are 9 months old in the US (Petts & Knoester, 2018). Other kinds of satisfaction also rose for new parents in Korea (Kramer et al., 2019). Paternity leave increased father's job satisfaction, which led to a rise in life satisfaction. The latter improved the mother's family relationship satisfaction.

This paper aims to analyze some of the effects that have been exposed in the previous literature review but with a focus at the long-run for the Spanish policy setting. The variables of interest are divided into four groups and reported separately in different subsections of the results part. These consist of labor market outcomes, fertility, satisfaction and childcare time. Finally, we will have a last line of study in which we will classify individuals into different groups, according to their wage gap, age and education gaps, and gender role views, so that we can observe whether the effects of the policy have been heterogeneous for different groups in the sample.

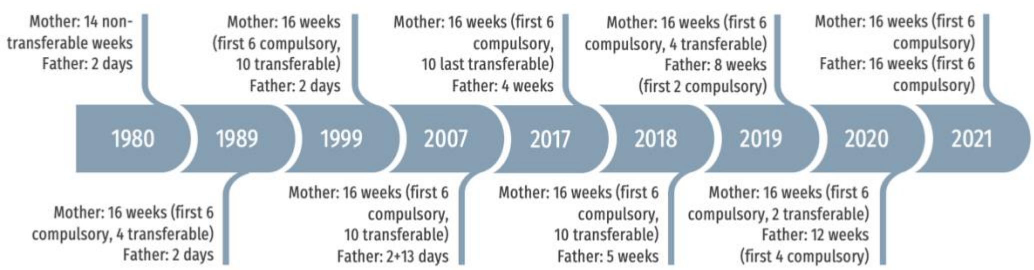
These issues are examined using Spanish Data from the Fertility Survey of 2018 runned by the National Statistical Institute (INE, 2018) through a regression discontinuity design (RDD) analysis of the responses of 931 and 960 parents, for the policy changes of 2007 and 2017 respectively. The remainder of this study is organized as follows. Chapter 2 presents the Spanish regulatory framework and societal situation necessary to understand the environment in which the policy was introduced. Chapter 3 consists of an outline of the data set used and a statistical description of the sample selected, as well as an explanation of self-created variables necessary for the analysis. Chapter 4 corresponds to the method used for the analysis. Chapter 5 stands for the results interpretation to answer the hypotheses from regressions. Finally, Chapter 6 contains the conclusions synthesizing the main points extracted, the limitations faced during the study and further research for future investigations.

2 The Spanish Institutional Setting

Spain is characterized by a lack of involvement in family policies, being one of the five OECD countries with the lowest public spending on it (OECD, 2019). Despite their invisibility, these policies have been gaining importance and attention from all political parties in recent years (Ayuso & Bascón, 2021). With respect to the conciliation of work and family life, exclusive paternity leave for fathers has undergone the greatest changes, significantly increasing its duration in the first two decades of this century. It was introduced as a gender policy to promote a more balanced share of family responsibilities, since the difference in the time that men and women spend on them is very large, and although it is getting smaller there is still a long way to go. According to the Spanish survey on time use from 2003 (INE, 2003), women spend 4 hours and 45 minutes compared to 2 hours and 8 minutes for men on these activities daily. The last data available from 2010 (INE, 2010) shows an equalizer trend, as men spent 2 hours and 37 minutes a day involved in activities related to family and care, whereas women spent 4 hours and 36 minutes.

In 1980, the first maternity leaves of 14 weeks and 2 days of paid leave for fathers were introduced. In 1999, mothers had 6 weeks of mandatory maternity leave and 10 weeks of parental leave, all with full pay, which could be partially or fully transferred to the father. However, just over 1% of maternity leave was reassigned to fathers in 2018 (Seguridad Social, 2018), a proportion that remained stable since 2005 (Escot et al., 2014). According to a survey on Family and Changing Gender Norms (International Social Survey Programme, 2012) out of the Spanish people surveyed that thought that paid leave should be available to parents, 30% of respondents thought that the leave should be split equally. Societies' values were not guiding the decisions taken by the families. That could be the case because the default option was that moms took the parental leave, which makes families more likely to follow it (Jachimowicz et al., 2019).

Figure 1: Parental leave reforms in Spain



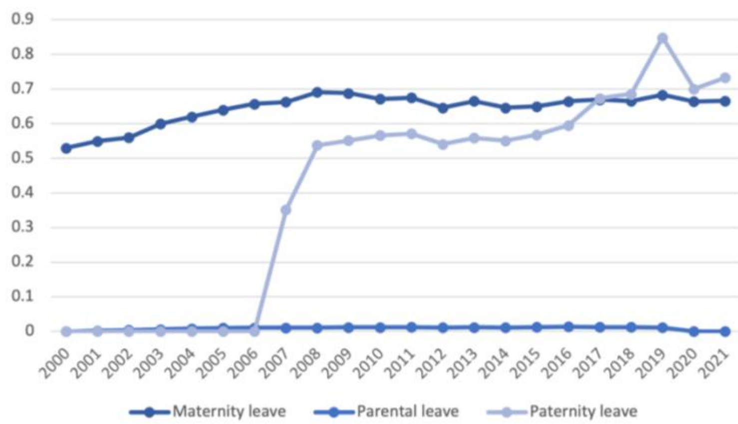
Evolution of paternity and maternity leave in Spain since its introduction
Source: Own elaboration

After the period of paid leave, both parents could enjoy unpaid leave of up to the time when the kid turns 3 years old, an individual and non-transferable right for all workers that allows them to return to the same job, the first year without affecting the accrual of the pension. Another option for one of the parents is to reduce the working day by up to 50% until the child reaches the age of 8, with a consequent reduction in salary. Data from the Spanish Labor Force Survey (INE, 2006) show that mothers worked more part-time because of family responsibilities (17% compared to 0.2% of men) and took more unpaid leave (3.8% of women compared to less than 1% of men).

As of March 24 from 2007, new fathers were eligible for a fully compensated 13-day paternity leave period, provided they were affiliated to Social Security and had worked for at least 180 days over the previous 7 years. Prior to the introduction of this law, there was little social and political debate on gender inequality and men's involvement in childcare. The reform took many families by surprise, making it difficult to take it into account when deciding when to have a child. Thanks to the implementation of the Spanish Equality Act (BOE, 2007) 13 days paid by the social security were added to the previously existing 2-day childbirth leave paid by the company, summing up total of 15 non-transferable days for fathers. In addition, some companies offered before and after the reform own financed extra days of paternity or maternity leave beyond the general labor law. Paternity leave was extended to four weeks in January 2017, despite its planned extension in 2011. These first two reforms are the ones studied in this article. Expansions of paternity leave followed every year one after the other (see Figure 1), until reaching the equalization of the leave time reserved for the mother in 2021. Out of the 16 weeks that fathers (or the member of the couple not giving birth to the baby) have nowadays, 6 have to be taken right after the birth. The last 10 weeks can be chopped up, posing a challenge for some companies (Pérez, 2023). However, this problem is not so exacerbated right now, as in 2021 76% of fathers took their leave continuously (PPiiNA, n.d.).

The last rise in the weeks of paternity leave positions Spain as the second country from the European Union with the longest available leave exclusively for fathers after Finland (European Parliament, 2023). As mentioned, the replacement income is 100% and the ceiling limit is generous (€3,642 a month in 2016), which by far exceeds the median income of the country at €1,929.70 (€1,677.60 for women and €2,160.40€ for men) (INE, 2016). Thus, most parents receive 100% of their income, being strong economic incentives to take both maternity and paternity leave. This has promoted a high take up, to the extent that since 2017 more paternity leave than maternity leave is requested in Spain (see Figure 2). In 2021, when the weeks of the paternity leave were equalized to the ones of the mothers, almost 75 percent of new dads took them up. Nevertheless, resistance to the use and different discourses according to ideology still exists (Barbeta-Viñas & Muntanyola-Saura, 2021; Romero-Balsas et al., 2013). Lastly, as already stated, before mothers and fathers had the same number of weeks available less than one percent of the fathers used the weeks that they could share with the mother. This is observable in the time trend of parental leave in Figure 2.

Figure 2: Fraction of mothers and fathers taking the leaves



Evolution up to the present day of the take-up of paid maternity and paternity leave, as well as of the numbers of shared weeks that the father could take from the mother. Computed dividing the number of permits of each kind requested in a given year by the total number of births of the same one.

Source: INE (annual number of births) and Seguridad Social (annual number of leave takers). *The value from 2019 is reported in two different data sets due to changes in the legislation. There seems to be issue with the reporting from the Spanish government due to an unexplained spike in the data.

3 Data

3.1 Sample description

The microdata set analyzed is the Fertility Survey (INE, 2018), conducted at the beginning of 2018 which includes parents who had kids around the two dates of interest, 2007 and 2017. By the time the data was obtained, the kids were between 10 and 12 years old for the 2007 threshold, while 0 to 2 for the 2017 one. The information was collected over a period of 15 weeks, from March 12 to June 25, 2018, for people between 18 and 55 years old. The INE obtained two independent samples for each of the 17 autonomous communities, one of men and a larger one of women. The type of sampling consisted of a stratified two-stage, which allows to obtain more precise estimates by considering the heterogeneity of the population through stratification. The questionnaires were completed via web, telephone interview and computer-assisted personal interview, concluding with a sample of 17,175 individuals (14,556 women and 2,619 men) distributed in 1,886 census sections. To have the complete data set used in this research, two data sets regarding two different parts of the questionnaire were merged, one with data from the interviewed and another with answers about a specific child.

For the purpose of this research, only those men and women who had kids around the two increases in paternity leave in 2007 or 2017 are the target. The observations from the person with the identification number 9193 are deleted because she had three kids around both thresholds and can be considered to receive the treatment of paternity leave expansion twice. For the specific sample used, five additional restrictions are imposed. First, same sex couples are excluded because the interest relies in gender specialization. Second, only one child is considered per birth, ruling out multiple births that generate correlated observations. Third, the observations corresponding to the couple's exclusive children are eliminated, since it has not been possible to enjoy the leave. Fourth, observations of people who had children with some previous couple are eliminated for two reasons. If they don’t have a new life partner many variables are not reported, and if they do it is not possible to know if the ex-couple or the actual partner is the parent of the kid studied. Finally, as paternity leave in the public sector has covered 10 days since January 2006 due to the introduction of “Plan Concilia”, I eliminate the cohort of fathers working in public positions. They are a big group, 17 percent of the observations, that do not enjoy such a huge increase in the numbers of days available. Additionally, birth order is controlled for, as parental investment has been found to differ accordingly, being higher from both parents the less kids they have at a given moment (Price, 2008).

For the cutoff of 2007 I end up with a sample of 147 men and 784 women for the main specification of a 1-year window, and similarly with 168 men and 792 women for the 2017 threshold. As it can easily be realized, despite collecting data from fathers and mothers, the sample of fathers is smaller and therefore it might not yield significant results independently. The number of father responses is smaller not because of a lower response rate or any other reason that might affect the validity of the data, but because of INE's decision on the sample size for each sex. No explanation is given for this choice, it is only mentioned that this is the first time that the fertility survey includes the responses of men. For some variables (like taking paid or unpaid leave, age, education, percentage of housework done…) data for both members of the couple per each observation is available, which somehow eliminates the problem of having so few men. The inconvenience that arises using this approach is that the perception of the division of domestic and care tasks for men and women is different, as men report higher percentages of participation in these tasks than their partners (Koster et al., 2022; Shafer et al., 2020). Hence, a potential issue that could arise analyzing the mother's and father's views together, for example regarding the housework done by each member of the couple, is that the results are closer to the mother's views as they are overrepresented.

The descriptive statistics of the two data samples from each time period are reported in Tables 1 and 2. In 2007, the average number of weeks of maternity leave taken by women was well below 16 weeks, while in 2017 the value was very close to the target. For men, the average value of weeks around April 2007 is above two weeks, when it should be below. This may be because many men used holiday days as part of their leave, as González & Zoabi (2021) find. For both samples, the average age of mothers is three years younger than that of fathers, with progenitors of children born in 2007 being older, as it is logical. However, in 2017 parents were older when they had their children, as the difference between their ages comparing the two data sets is only 6 years when it should be 10 if everything remained the same. Fertility changes are also reflected in mothers having their first child later, a variable that has increased in mean value by two years from 2007 to 2017, and in a drop in the number of children per person. In contrast to the birth rate, the average education of both fathers and mothers has increased between the ten years, with the level attained by mothers being the highest of both sexes in both samples. Finally, we also observe a drop in the number of married couples between the two time periods and a higher proportion of mothers working full time. All these trends have occurred in the Spanish society over the decade (Eurostat, 2023a, 2023c; INE, 2021), which allows us to observe a first resemblance between the data set and the population of interest.

3.2 Created variables

Two additional variables are constructed to be able to separate individuals into groups. First, inspired by González & Zoabi (2021), a pay gap measure is created using the net monthly income of each couple reported in intervals of 500 euros. The variable equals 0 if the individual has no income, 1 if it is below 500 euros, 2 if the salary is between 500 and 1000 euros etc. I subtract the mother from the father's interval and classify as egalitarian those couples with a pay gap of -1 or 0, intermediate those with values of 1 and 2, and high from 3 to 7. This measure is more exact to quantify the actual pay gap than the one González & Zoabi (2021) use. At their study, they use the differences in education and age to classify the couples according to their potential pay gaps. The authors find that the effect of the introduction of two weeks of paternity leave in 2007 generated changes mainly on couples classified as intermediate. In egalitarian couples, women did not specialize in childcare already before the policy change, and in high couples it was still beneficial for the family's economy for women to specialize even after the introduction of the policy.

Due to the distinction between the construction of our measurement and the one used by the researchers, I will also divide the groups according to education and age. These variables are less likely to be outcome variables than the pay gap, which can vary through the changes introduced by the policy. Age is impossible to modify and education takes longer and more effort to materialize than a wage rise or drop. Apart from the division between egalitarian, intermediate and high couples generated using the variable pay gap, the same three groups (egalitarian GZ, intermediate GZ and high GZ) will be created using age and education gaps (father's age or schooling minus mother's age or schooling) according to the values given by González & Zoabi (2021). The only difference relies in the fact that the education gap variable had to be adjusted since the data set does not provide education in years. Thus, the gaps of -2 and 4 years of education gap they use to construct the thresholds are equalized to -1 and 1 levels of education gap, which is not really precise.

The classification is displayed in Appendix Figure A1, resulting in the following thresholds:

- Egalitarian GZ: (i) age gap up to 1 year and education gap up to 1 level or (ii) age gap up to 3 years and education gap up to -1 level.
- Intermediate GZ: (i) age gap up to 1 year and education gap more than 1 level or (ii) age gap between 1 and 3 years and education gap between -1 and 1 levels.

- High GZ: (i) age gap more than one year and education gap more than 1 level or (ii) age gap more than 3 years.

Furthermore, 6 questions were selected to which the respondent could answer agree, neither agree nor disagree, or disagree. I assign a value of 0 to the most egalitarian and 2 to the most traditional view. If the person agrees with the following 4 questions, they will have a value of 2 for the given answer: "For a woman, the priority should be her family rather than her career"; "If a woman earns more money than her partner, this is not good for the relationship"; "When jobs are scarce, men should have more right to a job than women"; and "Taking care of the house and the family is as fulfilling as paid work". On the contrary, if the individual disagrees with the following 2 questions, they will have a value of 2 for the answer: "Men should participate in housework to the same extent as women"; and "A working mother can have as close a relationship with her child as a non-working mother". Once the responses to these 6 questions have been recoded, I perform the principal components analysis technique to find combinations that explain most of the total variability in the data, thus also eliminating problems of multicollinearity between responses that may be related. Finally, the values are re-scaled from 0 to 1 to make it easier to interpret the responses, with the value 0 corresponding to individuals with more equal values and beliefs. This variable will be referred to throughout the study as gender norm, the higher the value the higher the gender norms the respondent has. Separate regressions will be runned for those observations with a value of gender norm below and above the median of the variable, to see whether the results of the introduction and extension of paternity leave are different according to the values of the treated. The median is used as the partition point instead of the mean to have the same number of observations in both groups, as the distribution of the variable gender norm for both 2007 and 2017 is really right skewed, which means that the sample has more people with more equal values.

4 Methodology

The effect of the introduction of two weeks of paternity leave and the extension ten years later to four weeks will be studied in this paper. The research questions will be answered through a regression discontinuity approach, exploiting the differences in fathers' exposure to the paternity-leave quota depending on the month and year of birth of the child. One should not forget the importance of the word exposure, as the effect of the two-week paternity leave on those who took it is not studied, but rather the effect on the families who had the possibility to take it. The main variables of interest to study whether they vary around the cut-off point are absence for childcare taken by fathers, kids desired, childcare and housework responsibility, and the well-being of both parents. Two different cut offs will be used: the increase to 15 days in 2007 and the increase to a month in 2017, as both of them can be analyzed using the same data set. Nonetheless, the interest on long term effects of paternity leaves can only be studied for the 2007 reform, as the data was collected too close to the second cutoff date.

The regression discontinuity design is a strategy that allows to estimate the causal effect of a treatment, in this case the increase in the number of days of paternity leave. Individuals are split into two groups and compared, based on their position relative to a cut-off point, which in this case is either April 2007 or January 2017. The differences in outcomes, the dependent variables, between the two groups are compared to estimate the causal effect of the treatment. The effect is considered causal because the assignment to each side of the cut-off is non-random but abrupt, creating an exogenous source of variation. There are two premises needed for this to be true: that individuals close to the cutoff are similar in observable and unobservable characteristics, and that there is no bunching right before or after the cutoff, which would mean that individuals had some manipulation available. These two assumptions will be checked in the upcoming results section.

The estimated equation will be, for each period:

$$Y_{im} = \alpha + \beta T_{im} + \delta_1 m + \delta_1 T_{im} m + \gamma X_{im} + \epsilon_{im}$$

Where Y is the dependent variable of interest (e.g. satisfaction of mothers regarding childcare sharing) for family i who had a child in month t, T indicates paternity leave eligibility (being β the coefficient of interest), m is the month of birth of the child, and X are control variables. The paternity leaves were introduced on the 24^{th} of March and 1^{st} of January respectively. The nature of the data only provides information about the month of birth, so April 2007 and January 2017 are normalized to 0 (corresponding March 2007 and December 2016 to -1, May 2007 and February 2017 to 1 and so on). The month of birth and its interaction with being born after the policy introduction are included as controls in the regression, because the age of the kid or baby might have a significant effect on the outcome variables.

When the dependent variables are binaries, logistic regressions are performed, while when they are continuous linear regressions are conducted. Different windows around the threshold are included, from 12 months to 3 months before and after the cutoff date. During the chosen time periods, there should be no significant differences in the socioeconomical conditions the families were exposed to, as the economic recession hit Spain around September 2008. This is important to ensure that the effects of the paternity leave reform are not mixed with other factors that could interfere with the outcomes.

5 Results

The first step is to conduct the two validity checks for the RDD strategy selected. First, it is tested if the number of births around the threshold is continuous and therefore no bunching is found, which translates into families having no control on the group they want to belong to. Both policies from 2007 and 2017 were announced too short in advance, so parents could not decide to conceive at that time and select intro the treatment group right after the paternity leave expansions were introduced. Nevertheless, it would have been possible that families postponed a programmed induction on cesarian section to become eligible for longer paternity leaves. At a glance, no significant discontinuous jump is observed (Figures 3 and 4). The same possible behavior is tested formally applying equation (1), being the outcome variable the number of children born per month. Results are reported in Appendix Table A1 for the same three different specifications that will be used during the whole results section. For none of the time windows in both increases in paternity leave length studied the coefficients are statistically different for zero. These two results together show that families did not have any significant control to get or avoid the new policies.

The second required test is to check if the families having a baby before or after the threshold were balanced in their observable characteristics. Again, equation (1) is estimated but using mainly mother and father's characteristics as the outcome. Appendix Tables A2 and A3 report the results of these estimations for 2007 and 2017, respectively. In the names from the tables, the year 2007 or 2017 in the title corresponds to the birth cohort used, but it should not be forgotten that all data has been collected in 2018. Eighteen different characteristics are chosen, one regarding the infants (the sex of the newborn), seven for the mother and father separately (age, age when they had their first kid, education, three labor market characteristics and nationality), one that uses data from both parents simultaneously (paygap) and two measurements that correspond to a men or women depending on the observation (married and gender norm). The only two characteristics that are significantly different for the groups exposed to an expansion of two weeks in 2007 are father's age and marriage status, but only at the 10% significant level and for one of the studied thresholds. On the other hand, the cohort exposed in 2017 to an increase to 4 weeks of paternity leave had less educated mothers from other countries, being both of the coefficients statistically different from zero at the 90% confidence level. Hence, these variables will be added to each regression respectively, depending on which data regarding the 2007 or 2017 cohort is being used, to control for the significant differences.

Finally, it is important to remark that values for the variable gender norm are balanced around the threshold in 2007 but not in 2017. This translated into individuals not having more gender egalitarian views right after the paternity leave introduction 2007. This is important because it cannot be concluded that fathers who were eligible for longer paternity leaves are more involved in childcaring because their own values or the values of their wives have changed. The reason for this lack of increase in individuals having more egalitarian views may be due to the fact that the introduction of two weeks of paternity leave in Spain was established abruptly and without prior debate in society about it. The individuals affected by the policy may only have welcomed it but not reflected on their role in the face of greater gender equality. On the opposite hand, individuals who were eligible for 4 weeks of paternity leave in 2017 had more gender egalitarian views after the introduction of the treatment, and that might be a trigger for fathers becoming more active in childcare activities.

The two validity checks performed for each threshold support the RDD identifying assumptions, allowing this method to be valid for this study. Most controls are not included in the main analysis, because if the second assumption of RDD holds, observed and unobserved covariates are balanced and they don't need to be incorporated. This application allows the number of observations not to drop because of some missing values in the controls. Finally, another fact that allows to see that the values of the chosen data set resemble reality is that the number of fathers who took paid leave in the year after the introduction of the policy was very similar to the proportions observed in Figure 2. The takeup was 51.1% in 2007 and 68.5% in 2017, computed dividing the number of fathers who took the paternity leave by the births in that given year.

5.1 Labor market outcomes

More than ten years after the introduction of a paternity leave of two weeks in 2007, some differences are still found between the treated and untreated groups (see Table 3). Fathers' labor market outcomes differ, noting an increase in the likelihood of the father working but also in the number of part-time contracts, but only for one time window at the 5% significance level for both variables. Put together, the results suggest that men who would have decided not to work ended up working part-time instead. Ten years later, it may be considered common for this part-time working arrangement to continue, as such contracts often remain a long-term arrangement (Riederer & Berghammer, 2020).

Additionally, another effect on male behavior was a high increase in leave of absence for childcare. The coefficient only appears for the window of 12 months, because in the other two specifications the results are not reported due to perfect collinearity. This occurs because in the untreated group there are no fathers that took this kind of unpaid leave, and all the fathers who took it had kids after the threshold of April 2007. Therefore, the model predicts that the results will be perfectly forecasted depending on which group we are in. This problem arises due to the low number of unpaid parental leaves requested by fathers, around one percent for the Spanish population as a whole. Hence, there exists a high probability of the number being zero in a large part of the months of the sample used. This low number could also explain why Farré & González (2019) do not find a higher take-up of this leave after the policy change. They study if fathers took it 6, 12 and 24 months after, so at an exact moment in time. As only 1% of dads took the unpaid leave back then, the chances of finding fathers that were taking in it in a specific moment are way lower than studying if they ever took it during the 3 years possible. Leave of absence for childcare is an outcome that cannot be classified as long-term, as it can only be taken until the child is 3 years old, and hence the results would be the same 10 years after the birth or 7 years before it.

The effects that do appear to be significant for more than one window around the threshold, according to the results shown in Table 3, are the likelihood that the mother has a full time position, a decrease in the number of weeks of maternity leave, and an increase in the number of weeks of paternity leave for fathers. This last effect was expected since that was the aim to the policy, to increase the number of days of fathers' leave from 2 to 15. To compute the change in the number of weeks of paid maternity leave taken by women, robust estimators have been used as there are outliers in the data for mothers who took 80 and 99 weeks of leave, which is very unlikely to be given voluntarily by a private firm. This decision of going back to work sooner after the birth, together with the increased likelihood of mothers working full-time even ten years after the introduction of the leave, implies that there are spillover effects on the decisions taken by mothers, as less of the sole responsibility for childcare falls on them after the introduction of the policy.

On the contrary, the extension of paternity leave to four weeks in 2017 only led to significant short-term changes in the weeks of leave taken by fathers, but not on any other labor market outcome (result summarized in Table 4). Despite of that, the two coefficients regarding the leave of absence for childcare taken by fathers have again a positive sign. As previously stated, the results from both cutoffs cannot be compared, as in this last case de kids are only between 0 and 2 years of age while for the 2007 cutoff they are between 10 and 12. Additionally, the control group for the 2017 cutoff are families who had available 2 weeks of paternity leave.

5.2 Fertility outcomes

Another area of interest is the influence of paternity leave on long-term fertility. Total fertility has not been studied in the long term, only birth spacing between children after an introduction or increased paternity leave. Cools et al. (2015) conclude using data from Norway obtained 14 years after the reform that the fertility gap did not vary. In contrast, Farré & González (2019) continue to find 6 years after the introduction of the 2007 paternity leave in Spain a lower likelihood of having one more child. The birth spacing is found to be between 1 and 4% larger when taking into account all women, and between 2 and 9% larger for those women who had their first child after turning 30. Their conclusion is that as older women are closer to the end of their fertile cycle, this delay may have affected their full fertility.

When the total number of children 10 years later is studied with the actual sample, there are no significant differences between the decisions made by families exposed to the 15-day leave and those not exposed to it. The effect is not the same when only those observations in which the mother was over 30 years old when she had her first child are selected. In this group, which is the one that Farré & González (2019) find to increase the time between the birth of their kids, women have between 0.287 and 0.439 more children if their partners were exposed to the choice of taking paternity leave (specifications shown in Table 5). These values are significant and high in magnitude, representing an increase over the average number of children of between 14.7 and 22.2%. This implies that access to paid leave increases fertility intentions at the intensive margin. This may be due to the fact that the mean number of children for women who have their first child before 30 is 2.41 while for those who have their first child after 30 it is 1.96. The average number of children desired by women in each group is 2.62 and 2.45 respectively. Thus, it is the women who had their first child over 30 who were farthest away from their desired fertility, and thus the policy is likely to have had a greater effect on their total fertility.

One of the most different variables in our data set is the number of desired children, which has never been studied in relation to paternity leave in either the short or long term. While the number of children desired by women is not significant and also inconclusive in terms of whether there is a positive or negative change with the introduction of leave, I find that men who had children from April 2007 have positive coefficients for this variable for all of the three time windows despite not being significant. Therefore, it can be concluded that fertility preferences did not decrease as Farré & González (2019) insinuate. In their research, they compare the number of children desired by fathers and mothers in 2001, 2006 and 2011, concluding that men reported significantly lower desired fertility in 2011 relative to women while in 2006 it was the other way around. Although they cite that there may be other explanations, they explain that men's greater participation in childcare, driven by the introduction of paternity leave in 2007, could have caused this. However, after analyzing desired children through RDD, the fall in desired children that the researchers find for men would be attributed to other factors rather than to their greater involvement in child rearing. One of the main reasons that may have caused men to go from wanting more children than women in 2006 to fewer in 2011 could be the unequal way in which the 2008 economic crisis unequally affected both sexes in Spain. Unemployment rose during that year by 78.1% for men and 24.7% for women, due to the fact that the most affected sectors, such as construction, were predominantly male dominated. For the first time in the country's history, the total number of unemployed women was lower than that of men (El Mundo, 2009).

Again, the expansion of paternity leave in 2017 had virtually no effect on actual or desired fertility. In Table 6 it is observable that the treatment did not generate any causal effect. Nonetheless, as for the previous reform of 2007 all the coefficients of the desired number of kids by fathers are found to be positive.

5.3 Childcare time

As it has already been shown several times through data, it is mothers who spend the most time on care activities. This is why, in this subsection, I create binary variables for ten different care activities. The time spent on childcare performing different actions needed when raising a kid can be interpreted as the parental investment that fathers dedicate to their children. In the survey, the interviewee is asked questions about who is in charge of carrying out various activities such as dressing the children, deciding their food etc. To these, one can answer: oneself, one's partner, both, or other options such as the child him/herself or the grandparents. The latter are discarded, as they are not relevant to what is being studying. I construct a variable for each activity, which takes a value of 0 if it is the mother who oversees the activity and 1 if it is both or only the father. These two options are put together since in very few cases mothers do not participate in an activity at all. Thus, a positive coefficient will mean that fathers in the group that was affected by the reform participate more in care activities, either by themselves or sharing the tasks with the mothers.

As Table 7 shows, all coefficients are positive, which translates into a significant increase in the likelihood of fathers being more involved in numerous caregiving tasks even 10 years later. The participation rose significantly for buying clothes, play, take care when sick, help with homework and choose the food. A more equal participation in the activities of dressing the kids and choosing extracurricular activities is also found but only for one time window at the 10 percent significance level. It is interesting to highlight that the paternity leave introduction of two weeks rose the participation of fathers in the four caring activities that were more unequally distributed in the sample prior to April 2007, which are dressing the kids, choose the food, take care when sick and buy them clothes (the mean of the distribution of this activities can be found in Appendix Table 11). An explanation for not finding any or hardly any effects on bathing, dressing and laying kids down could be that, as the children studied are 10 years old, there are far fewer observations (almost half) on these variable as at this age they are often already doing them independently.

In the case of 2017, less coefficients are significant compared to the ones from the 2007 reform. Expanding paternity leave to four weeks mainly increased the likelihood of fathers putting children to bed and playing with them (these results can be found in Table 8). Fathers also were more likely to participate in the activities of dressing the baby and choose extracurricular activities, but the effect was only found significant for one time window at the ten or five percent significance level.

In addition to the fact that much smaller effects are found after the extension of the permit in 2017 than after its creation 10 years earlier for all the groups of variables studied, two possible explanations may exist for the less pronounced increase in the participation of fathers in childcare activities. On the one hand, the averages of the participation in the families that had kids prior to the expansion of the leave to four weeks show a much greater parental involvement from fathers in all care activities (Appendix Tables A4 and A5). In addition to the fact that the "untreated" comparison group for the policy change in 2017 was being treated with two weeks of paternity leave and not two days as in the 2007 case, social norms and fathers' involvement in caring tasks have become more egalitarian over the last few years in Spanish society (see data from Chapter 2). Another reason for finding smaller effects could be that the families studied in this case have younger children, aged between two years and months. As Borràs et al. (2021) explained by after a study in Spain, the younger the children are, the more mothers are responsible for their care and household activities. The latter trend also appears in the 2007 data for household activities carried out by fathers and mothers, with a more equal sharing as children grow older.

5.4 Satisfaction and implication in housework

The change in fathers' satisfaction on various issues is an understudied outcome following the introduction of paternity leave. Only live, work, and work-life balance satisfaction were previously studied in other European countries (Korsgren & van Lent, 2022), also in the long term. The authors could not explain where the rise in life satisfaction, the only one found, was coming from. The increase was especially high for the mothers. Using data for Spain 10 years after the introduction of the policy, it is observed that satisfaction with caregiving and with the relationship increase for the mother, being the first one especially significant and strong in magnitude (see Table 9 for the regression coefficients). The level of satisfaction regarding childcare sharing is measured from 0 to 10, being an increase between 0.88 and 1.45 high, specially taking into account that the questionnaire is answered by a time the kids were around 10 years old. Despite not being a significant reduction, all the coefficients of the regressions computed using father's satisfaction in childcare sharing are negative. Therefore, the increase in fathers' involvement in childcare activities shown in the previous subsection has led to a decrease in their satisfaction regarding this topic. In addition, men's satisfaction in house chores sharing and with his partner decreased but being the effect only significant for the 12 month specification and at the 10% significance level.

Finally, the percentage of housework done by each household member decreases for mothers and increases for fathers by a similar percentage. This would mean that the extra time that fathers enjoy being home after January 2007 translated in also a longer time expend doing house chores, even 10 years later. However, what might seem odd is that as a consequence, satisfaction regarding the division of housework did not increase significantly for mothers. One explanation is that the differences in the percentage of tasks performed by fathers and mothers, although reduced, are still very unevenly distributed. The averages of the percentage of tasks performed for the 3-month regression around the implementation of the leave are 65.6% for mothers and 27.5% for fathers. Therefore, in spite of an improvement, it is normal that mothers are still not satisfied with the distribution.

The change in the satisfaction with childcare division for 2017, as can be seen in Table 10, maintains the same trend as before. Only one of the coefficients, the decrease in father's satisfaction regarding the distribution of childcare between the couple, is significant at the 10% for one of the thresholds. The effect might not be high because childcare time by fathers did not change for six out of the ten activities studied, as exposed in the previous subsection.

5.5 Outcomes per groups

Lastly, the results of the variables that show an interesting trend in the sign or significantly varied with the introduction of the paternity leave reforms have been calculated independently for different groups of individuals, following the three possible divisions explained in Chapter 3. It should not be forgotten that the splitting mechanisms based on paygap and gendernorm are exploratory, as despite neither of them varying around the introduction of the policy in April 2007, they can be considered outcome variables. For example, chances exist that the variable paygap was affected if mothers increased the likelihood of being full time workers. In the case of the sample using births around January 2017, the interpretations should be done with even more caution as gendernorm does decrease significantly after the paternity leave expansion for a time window (Appendix Table A3). On the contrary, education and age generate heterogeneity in a predetermined way, so a third classification arises using the values provided by González & Zoabi (2021).

Before diving into the analysis of the significant coefficients, there are some interesting analyses of the observations. If the total number of observations of the year 2007 is compared with those of the year 2017 for the paygap-based specification (columns 1, 2 and 3 of Tables A6 and A7 in the Appendix), it is observable that the observations of the number of couples classified as egalitarian and intermediate increase, while those classified as high decrease. This coincides with the time trend present in the last decade in the Spanish society (Eurostat, 2023b). Secondly, if we look at the number of observations of women and men independently (those reported in the section of the variables regarding satisfaction) for the division based on gender norms (columns 7 and 8), we notice that for both time periods the same is true. Men have a higher number of respondents grouped as high gender norms (above the median value), while women have a higher number of respondents under low gender norms. This can be translated as meaning that at first glance, there are more women with egalitarian views than men. Nevertheless, we must be careful as due to the nature of the data set used there are far fewer responses from men than from women (and the sample becomes even smaller by dividing the men into groups), which may affect the reliability of the results.

With respect to the coefficients obtained from the RDD regressions, no very clear conclusion is obtained with respect to the divisions between egalitarian, intermediate or high. The interpretation of the results would be that, for example, the increase in fathers' involvement in the activity of helping children with homework or mothers' satisfaction in childcare is driven by couples with an egalitarian pay gap. Thus, a significant increase in childcare activity from fathers who have an intermediate or high pay gap with their partner is not found (see Appendix Table 6). It is not detected that the effect of the two-week paternity leave policy emerges mainly from intermediate couples ten years later, neither with González & Zoabi's (2021) specification (education and age gap) nor with the new one (paygap). It should be considered that our sample has far fewer observations, which may affect the reliability of the regressions when further splitting the data into smaller groups. The same inconclusive effects arise with the expansion of leave to four weeks in January 2017 (coefficients displayed in Appendix Table A7).

Finally, an interesting pattern is found when analyzing the outcomes per group divided with respect to the value of gender norm for the 2007 cohort. The introduction of the two-week paternity leave had significant effects in those families in which the respondent had a value of the gender norm variable below the median. Mothers from this group work more full-time 10 years later, had more children if the first child was born after the age of 30, and report more satisfaction in the sharing of care activities, household chores and with the partner. On the other hand, fathers are more involved in helping kids with homework and choosing extracurricular activities, and are less satisfied with the sharing of care and with the couple's relationship (column 7 of Table A6 from the Appendix). However, an increase in the gender norm variable decreases the likelihood that the father of the family took paid paternity leave, and it is thus logical that the effect of the policy is larger for this group. It should not be forgotten that, as this variable has been measured 10 years later and not at the time of policy implementation, the value of gender norm might also have changed as a result of taking the paternity leave.

6 Conclusion

Paternity leave has been at the center of family reconciliation policies in recent years. This has led numerous researchers to study their effects in different countries. Nevertheless, since they are fairly recent there is a lack in the literature of their long-term effects. In this dissertation, we have analyzed variables never observed in the long run after the introduction of paternity leave such as the number of desired children; the involvement of fathers in many different care activities; the satisfaction regarding childcare, housework or the couple; and the percentage of household chores performed.

The long-term effects found because of the introduction of paternity leave in 2007 are numerous. The hypothesis based on the ideas of Becker (1991) and Becker et al. (1985) is fulfilled. A short period of paternity leave has affected the evolution of household roles in the Spanish households. After dividing the results between labor market outcomes, fertility decisions and desires, childcare sharing and satisfaction, effects are found in all of them. On the one hand, spillover effects on the decisions taken by mothers are observed. Mothers of children born after the introduction of the policy have a higher likelihood of having a full-time position and took fewer weeks of maternity leave. Also, those women who started having children later had more. Third, there is a significant increase in the likelihood of fathers being more involved in numerous caregiving tasks, in addition to an increase in the percentage of household chores executed. This is accompanied by a significant increase in the mother's satisfaction with the repartition of caregiving. All coefficients regarding her satisfaction with her sentimental relationship and the sharing of household tasks are also positive, and the coefficients on fathers' reported satisfaction with all three aspects negative. Finally, the effects of the introduction of the two-week paternity leave are driven by those families in which the respondent had a value of the gender norm variable below the median. Nonetheless, as those individuals with a more egalitarian view ten years later are the ones who used the policy, we do not know whether taking paternity leave changed their gender norms, or whether it is the initial gender norms that affect both taking the paternity leave and families changing their behavior.

On the other hand, when applying RDD to the 2017 sample, the initial hypothesis is fulfilled. Only fathers' involvement in four out of ten caregiving activities significantly increases and their satisfaction with the division of the childcare tasks falls. The effects found in the short term, only one year after the extension of the paternity leave to four weeks in 2017, are smaller as anticipated but more non-existent than expected. This could be because parents in this sample had lower gender norms and fathers were more involved in childcare already before the expansion. Another possibility is that too little time has passed to observe changes, as the younger the child, the more the mother is involved in childcare (Borràs et al., 2021). It should not be forgotten that the results of 2007 are not comparable to those of 2017, as despite evaluating the same variables, they are not measuring the same thing. For the first time period long term effects are examined while for the latter they are short term, and the population studied has evolved over the ten years between the two policies.

One limitation of this study lies in the size representativeness of the sample. On the one hand, the use of data from very specific periods and from couples with certain characteristics (heterosexual, non-divorced, etc.) reduces the number of observations, making the estimates less accurate. On the other hand, the low percentage of men causes variables such as the desired number of children or men's satisfaction to have very few observations. It would also have been desirable to use bandwidths that include fewer months, but due to the low number of observations, the desirable adjustment between bias and variance prevents this from occurring with the available sample. The further away we are from the threshold, the more bias we introduce because there may be other factors impacting the outcome variable than the changes in the policy. Nevertheless, because the sample is so small, if we decrease the number of observations even further, the model will pay too much attention to each observation and will not generalize the results.

It can be concluded that paternity leave has been a beneficial measure for the Spanish society, at least until the studied reform of 2017. Neither its introduction nor its expansion has been found to have significant negative effects. An interesting reduction in fathers' satisfaction in several aspects like childcare sharing arose, which probably is driven by their greater involvement in childcare activities. This could affect their general well-being, but it would not necessarily be negative for society if it reduces the cost of childbearing for women. Thus, paternity leave does not make a choice in the trade-off between gender equality and parental investments in children, as maternity leave was generating when introduced alone. Only two weeks of paternity leave triggered larger changes in the within-household distribution of market and household work 10 years later. In the coming years, the effects of the extensions of paternity leave since 2017 until the equalization of the leave for both parents in 2021 need to be further studied. It is currently unknown whether these expansions have generated benefits for families and for gender equality, both in the short and long run. Their effect may have been attenuated, as the limited results found after the 2017 extension might suggest, by changing social norms to more gender egalitarian views over the last decade. Despite this possible development, it is the evolution of gender preferences and norms that are necessary to decrease gender inequality, but public policies must come along with this trend to accelerate change.

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Tables and Figures

Table 1: Descriptive statistics 2007

Variables	Number of observations	Mean	Standard deviation	Min	Max
Maternity leave length (weeks)	526	14.589	5.659	1	80
Paternity leave length (weeks)	334	2.476	2.704	1	19
Mother’s age	888	42.314	4.924	24	58
Father’s age	791	45.030	5.188	21	62
Mother’s education	931	1.177	0.667	0	2
Father’s education	914	0.977	0.696	0	2
Mother’s nationality	865	0.941	0.236	0	1
Father’s nationality	856	0.948	0.221	0	1
Mother working	654	0.917	0.275	0	1
Father working	895	0.887	0.317	0	1
Mother permanent contract	465	0.789	0.408	0	1
Father permanent contract	565	0.832	0.374	0	1
Mother full time worker	591	0.673	0.469	0	1
Father full time worker	615	0.951	0.215	0	1
Mother public worker	591	0.230	0.421	0	1
Father public worker	806	0.237	0.425	0	1
Number of children	931	2.161	0.834	1	8
Mother’s age at first child’s birth	924	29.024	5.126	14	45
Married	931	0.878	0.328	0	1

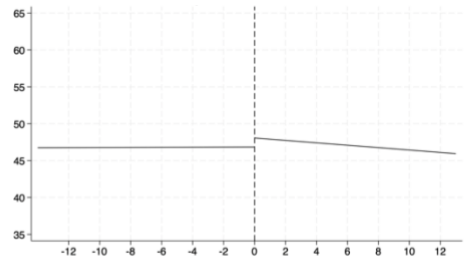
Note: The data comes from the Fertility Survey of 2018. The sample includes heterosexual fathers and mothers who had a kid with their current partner one year before and after April 2007. The variable of father public worker has been computed with the complete sample before deleting it, while the rest of the variables are reported without it as it is the main specification we use.

Table 2: Descriptive statistics 2017

Variables	Number of observations	Mean	Standard deviation	Min	Max
Maternity leave length (weeks)	642	15.718	3.742	1	40
Paternity leave length (weeks)	649	3.316	2.257	1	30
Mother’s age	931	34.896	4.901	18	51
Father’s age	807	37.627	5.483	19	60
Mother’s education	958	1.356	0.638	0	2
Father’s education	930	1.182	0.647	0	2
Mother’s nationality	859	0.950	0.218	0	1
Father’s nationality	843	0.943	0.232	0	1
Mother working	666	0.910	0.287	0	1
Father working	912	0.924	0.265	0	1
Mother permanent contract	486	0.775	0.418	0	1
Father permanent contract	669	0.770	0.421	0	1
Mother full time worker	598	0.711	0.454	0	1
Father full time worker	707	0.927	0.267	0	1
Mother public worker	598	0.271	0.445	0	1
Father public worker	707	0.222	0.415	0	1
Number of children	960	1.669	0.789	1	7
Mother’s age at first child’s birth	958	31.132	5.482	14	52
Married	960	0.714	0.452	0	1

Note: The data comes from the Fertility Survey of 2018. The sample includes heterosexual fathers and mothers who had a kid with their current partner one year before and after January 2017.

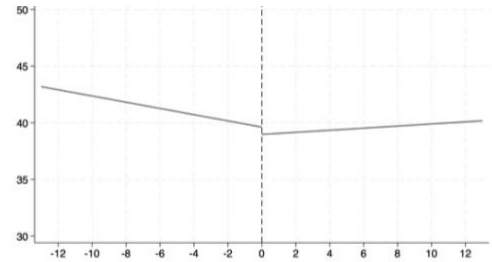
Figure 3: Children born per month around April 2007



Testing graphically for possible bunching of births around April of 2007 in the sample.

Source: Own elaboration with data set from INE (Fertility Survey of 2018).

Figure 4: Children born per month around January 2017



Testing graphically for possible bunching of births around January of 2017 in the sample.

Source: Own elaboration with data set from INE (Fertility Survey of 2018).

Table 3: Labor market outcomes 2007

Variables	(1) +/- 12 months	(2) +/- 6 months	(3) +/- 3 months
Mother working	-0.696	-1.033	-2.568
	(0.534)	(0.833)	(1.633)
Father working	1.283**	0.538	-1.519
	(0.596)	(0.750)	(1.291)
Mother full-time worker	0.146*	0.254**	0.400**
	(0.082)	(0.116)	(0.183)
Father full-time worker	0.077	0.085	-0.159**
	(0.040)	(0.051)	(0.072)
Mother's income	0.020	0.056	0.098
	(0.243)	(0.347)	(0.534)
Father's income	0.358	-0.074	-0.216
	(0.220)	(0.307)	(0.492)
Mother's reduced working hours	0.297	0.493	1.948
	(0.493)	(0.745)	(1.515)
Father's reduced working hours	0.710	2.905	
	(1.443)	(3.107)	
Unpaid maternity leave	-0.140	-0.273	-1.179
	(0.600)	(0.924)	(1.472)
Unpaid paternity leave	5.101***		
	(1.259)		
Weeks of mother's leave	0.179	-2.501**	-4.278***
	(0.944)	(1.248)	(1.604)
Weeks of father's leave	0.702	1.765**	1.662**
	(0.649)	(0.679)	(0.696)
N	931	459	236

Note: The data comes from the Fertility Survey of 2018. The sample used includes all heterosexual (non-public worker) fathers and mothers who had a kid with their actual partner in a specific time window of months (indicated in the column headers) around April 2007. The coefficients are from a dummy variable, the main independent variable, which equals 1 if the child was born after the introduction of the reform and therefore eligible for two weeks of paternity leave. The dependent variables (those related with the labour market) are the row header, and each coefficient is computed using a different regression (standard errors in parentheses). All the regressions control for a linear trend in the running variable (the month of birth centered at 0 in April 2007) and enable different trends before and after the reform date. Additional controls are the age of the father, being married and the order number of the child in the child table. The number of observations is reported for the specific time window, varying slightly for each variable due to missing observations.

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

Table 4: Labor market outcomes 2017

Variables	(1) +/- 12 months	(2) +/- 6 months	(3) +/- 3 months
Mother working	0.288	1.481	-0.067
	(0.656)	(0.908)	(1.320)
Father working	-0.209	-0.245	-0.368
	(0.601)	(0.881)	(1.364)
Mother full-time worker	-0.118	-0.096	-0.109
	(0.076)	(0.107)	(0.162)
Father full-time worker	-0.008	-0.010	0.018
	(0.042)	(0.068)	(0.101)
Mother’s income	-0.099	0.300	-0.486
	(0.200)	(0.292)	(0.430)
Father’s income	0.002	0.005	-0.003
	(0.192)	(0.271)	(0.391)
Mother's reduced working hours	0.643	0.913	1.222
	(0.473)	(0.705)	(1.142)
Father's reduced working hours	0.664	1.020	0.970
	(0.916)	(1.368)	(2.172)
Unpaid maternity leave	0.085	-0.368	-0.428
	(0.402)	(0.548)	(0.842)
Unpaid paternity leave	0.418	1.761	
	(1.008)	(2.476)	
Weeks of mother's leave	0.258	0.063	1.558
	(0.597)	(0.788)	(1.186)
Weeks of father's leave	1.878***	1.754***	2.006***
	(0.353)	(0.388)	(0.576)
N	960	488	247

Note: The data comes from the Fertility Survey of 2018. The sample used includes all heterosexual fathers and mothers who had a kid with their actual partner in a specific time window of months (indicated in the column headers) around January 2017. The coefficients are from a dummy variable, the main independent variable, which equals 1 if the child was born after the introduction of the reform and therefore eligible for four weeks of paternity leave. The dependent variables (those related with the labour market) are the row header, and each coefficient is computed using a different regression (standard errors in parentheses). All the regressions control for a linear trend in the running variable (the month of birth centered at 0 in January 2017) and enable different trends before and after the reform date. Additional controls are the mother’s education, mother’s nationality, and the order number of the child in the child table. The number of observations is reported for the specific time window, varying slightly for each variable due to missing observations.

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

Table 5: Fertility outcomes 2007

Variables	(1) +/- 12 months	(2) +/- 6 months	(3) +/- 3 months
Number of children	0.160	0.139	0.122
	(0.120)	(0.143)	(0.240)
N	931	459	236
Number of children people who had first kid older than 30	0.336***	0.287*	0.439*
	(0.115)	(0.159)	(0.263)
N	474	246	119
Desired children by woman	-0.058	0.043	-0.195
	(0.124)	(0.160)	(0.222)
N	784	379	195
Desired children by men	0.661	0.989	1.835
	(0.425)	(0.757)	(1.685)
N	147	80	41

Note: The data comes from the Fertility Survey of 2018. The sample used includes heterosexual (non-public worker) fathers and mothers who had a kid with their actual partner in a specific time window of months (indicated in the column headers) around April 2007. The coefficients are from a dummy variable, the main independent variable, which equals 1 if the child was born after the introduction of the reform and therefore eligible for two weeks of paternity leave. The dependent variables (those related with fertility) are the row header, and each coefficient is computed using a different regression (standard errors in parentheses). All the regressions control for a linear trend in the running variable (the month of birth centered at 0 in April 2007) and enable different trends before and after the reform date. Additional controls are the age of the father, being married and the order number of the child in the child table. The number of observations is reported for each variable for the specific time window, as the group of people used varies (in order: everyone, everyone if the mother had the first child when she was older than 30, only female respondents and only male respondents).

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

Table 6: Fertility outcomes 2017

Variables	(1) +/- 12 months	(2) +/- 6 months	(3) +/- 3 months
Number of children	0.097	0.101	0.164
	(0.098)	(0.134)	(0.211)
N	960	488	247
Number of children people who had first kid older than 30	-0.019	-0.131	-0.110
	(0.103)	(0.144)	(0.211)
N	643	323	155
Desired children by woman	-0.000	0.031	0.360
	(0.108)	(0.157)	(0.229)
N	792	400	204
Desired children by men	0.370	0.400	0.697
	(0.312)	(0.361)	(0.699)
N	168	88	43

Note: The data comes from the Fertility Survey of 2018. The sample used includes heterosexual fathers and mothers who had a kid with their actual partner in a specific time window of months (indicated in the column headers) around January 2017. The coefficients are from a dummy variable, the main independent variable, which equals 1 if the child was born after the introduction of the reform and therefore eligible for four weeks of paternity leave. The dependent variables (those related with fertility) are the row header, and each coefficient is computed using a different regression (standard errors in parentheses). All the regressions control for a linear trend in the running variable (the month of birth centered at 0 in January 2017) and enable different trends before and after the reform date. Additional controls are the mother's education, mother's nationality, and the order number of the child in the child table. The number of observations is reported for each variable for the specific time window, as the group of people used varies (in order: everyone, everyone if the mother had the first child when she was older than 30, only female respondents and only male respondents).

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

Table 7: Childcare time 2007

Variables	(2) +/- 12 months	(3) +/- 6 months	(4) +/- 3 months
Dress	0.323	0.453	1.591*
	(0.415)	(0.595)	(0.911)
Buying clothes	0.640**	0.694	1.373**
	(0.304)	(0.435)	(0.692)
Play	0.838**	0.369	1.035**
	(0.390)	(0.549)	(0.812)
Take care when sick	0.639**	0.691*	1.340*
	(0.303)	(0.435)	(0.690)
Help with homework	0.732**	0.905**	1.047*
	(0.303)	(0.437)	(0.665)
Choose the food	0.196	0.523*	1.464**
	(0.309)	(0.440)	(0.703)
Take to school	0.065	0.050	0.842
	(0.303)	(0.431)	(0.661)
Bathe	0.229	0.453	0.646
	(0.418)	(0.595)	(0.900)
Lay down	0.044	0.082	1.176
	(0.373)	(0.534)	(0.844)
Choose extracurricular activities	0.261	0.471	1.241*
	(0.340)	(0.486)	(0.728)
N	931	459	236

Note: The data comes from the Fertility Survey of 2018. The sample used includes all heterosexual (non-public worker) fathers and mothers who had a kid with their actual partner in a specific time window of months (indicated in the column headers) around April 2007. The coefficients are from a dummy variable, the main independent variable, which equals 1 if the child was born after the introduction of the reform and therefore eligible for two weeks of paternity leave. The dependent variables (those related with fathers' investment in childcare time) are the row header, and each coefficient is computed using a different regression (standard errors in parentheses). All the regressions control for a linear trend in the running variable (the month of birth centered at 0 in April 2007) and enable different trends before and after the reform date. Additional controls are the age of the father, being married and the order number of the child in the child table. The number of observations is reported for the specific time window, varying severely for each variable due to the different involvement of parents in different activities as children are to a certain extent self-sufficient.

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

Table 8: Childcare time 2017

Variables	(1) +/- 12 months	(2) +/- 6 months	(3) +/- 3 months
Dress	0.461	0.881**	0.046
	(0.284)	(0.421)	(0.651)
Buying clothes	0.021	0.543	0.025
	(0.285)	(0.418)	(0.644)
Play	0.770**	1.504***	1.859**
	(0.355)	(0.541)	(0.881)
Take care when sick	0.071	0.486	0.667
	(0.285)	(0.420)	(0.651)
Help with homework	0.053	0.559	0.277
	(0.297)	(0.431)	(0.683)
Choose the food	-0.214	0.461	0.747
	(0.296)	(0.441)	(0.694)
Take to school	0.020	0.148	-0.629
	(0.295)	(0.425)	(0.665)
Bathe	-0.068	0.285	0.563
	(0.292)	(0.425)	(0.655)
Lay down	0.463	1.020**	1.326**
	(0.289)	(0.425)	(0.654)
Choose extracurricular activities	-0.123	0.636	1.461*
	(0.333)	(0.492)	(0.805)
N	960	488	247

Note: The data comes from the Fertility Survey of 2018. The sample used includes all heterosexual fathers and mothers who had a kid with their actual partner in a specific time window of months (indicated in the column headers) around January 2017. The coefficients are from a dummy variable, the main independent variable, which equals 1 if the child was born after the introduction of the reform and therefore eligible for four weeks of paternity leave. The dependent variables (those related with fathers' investment in childcare time) are the row header, and each coefficient is computed using a different regression (standard errors in parentheses). All the regressions control for a linear trend in the running variable (the month of birth centered at 0 in January 2017) and enable different trends before and after the reform date. Additional controls are the mother's education, mother's nationality, and the order number of the child in the child table. The number of observations is reported for the specific time window, varying slightly for each variable due to missing observations.

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

Table 9: Satisfaction and housework 2007

Variables	(1) +/- 12 months	(2) +/- 6 months	(3) +/- 3 months
Satisfaction in childcare sharing mother	0.884**	1.152**	1.454*
	(0.387)	(0.569)	(0.882)
Satisfaction in house chores sharing mother	0.525	0.383	0.494
	(0.435)	(0.639)	(1.005)
Satisfaction relationship mom	0.264*	0.354	0.466
	(0.230)	(0.336)	(0.482)
N	784	379	195
Satisfaction in childcare sharing father	-0.322	-0.723	-1.797
	(0.519)	(0.809)	(1.200)
Satisfaction in house chores sharing father	-0.740	-0.681	-2.267*
	(0.554)	(0.824)	(1.247)
Satisfaction relationship father	-0.760*	-0.361	-0.810
	(0.399)	(0.558)	(0.994)
N	147	80	41
Percentage of housework done by mother	-3.006	2.407	-8.173*
	(3.022)	(4.331)	(6.732)
Percentage of housework done by father	3.396	2.538	11.540**
	(2.668)	(3.667)	(5.737)
N	961	459	236

Note: The data comes from the Fertility Survey of 2018. The sample used includes all heterosexual (non-public worker) fathers and mothers who had a kid with their actual partner in a specific time window of months (indicated in the column headers) around April 2007. The coefficients are from a dummy variable, the main independent variable, which equals 1 if the child was born after the introduction of the reform and therefore eligible for two weeks of paternity leave. The dependent variables (those related with satisfaction and housework) are the row header, and each coefficient is computed using a different regression (standard errors in parentheses). All the regressions control for a linear trend in the running variable (the month of birth centered at 0 in April 2007) and enable different trends before and after the reform date. Additional controls are the age of the father, being married and the order number of the child in the child table. The number of observations is reported per pools of variables for the specific time window, as the group of people used varies (in order: only female respondents, only male respondents, and everyone).

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

Table 10: Satisfaction and housework 2017

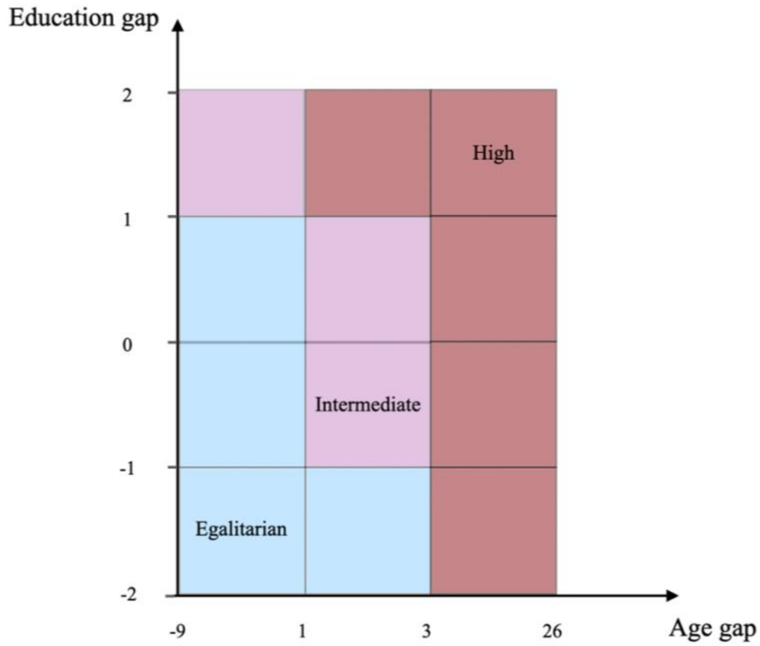
Variables	(1) +/- 12 months	(2) +/- 6 months	(3) +/- 3 months
Satisfaction in childcare sharing mother	0.043	0.291	0.106
	(0.346)	(0.511)	(0.733)
Satisfaction in house chores sharing mother	0.053	0.139	-0.035
	(0.382)	(0.549)	(0.807)
Satisfaction relationship mom	0.138	-0.074	-0.124
	(0.209)	(0.291)	(0.395)
N	792	400	204
Satisfaction in childcare sharing father	-0.507	-0.512	-2.376*
	(0.620)	(0.773)	(1.370)
Satisfaction in house chores sharing father	-0.176	0.117	0.318
	(0.562)	(0.770)	(1.171)
Satisfaction relationship father	-0.073	-0.294	0.090
	(0.374)	(0.633)	(1.132)
N	168	88	43
Percentage of housework done by mother	-0.027	-5.654	-5.758
	(2.630)	(3.924)	(6.947)
Percentage of housework done by father	1.370	5.797	7.706
	(2.367)	(3.569)	(5.413)
N	960	488	247

Note: The data comes from the Fertility Survey of 2018. The sample used includes all heterosexual fathers and mothers who had a kid with their actual partner in a specific time window of months (indicated in the column headers) around January 2017. The coefficients are from a dummy variable, the main independent variable, which equals 1 if the child was born after the introduction of the reform and therefore eligible for four weeks of paternity leave. The dependent variables (those related with satisfaction and housework) are the row header, and each coefficient is computed using a different regression (standard errors in parentheses). All the regressions control for a linear trend in the running variable (the month of birth centered at 0 in January 2017) and enable different trends before and after the reform date. Additional controls are the mother's education, mother's nationality, and the order number of the child in the child table. The number of observations is reported per pools of variables for the specific time window, as the group of people used varies (in order: only female respondents, only male respondents, and everyone).

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

Appendix

Figure A1: Classification of couples according to their differences in education and age



Classification according to the wage gap following the study and values of González & Zoabi (2021). The gaps are computed subtracting from father's age or schooling the mother's age or schooling. The thresholds for age are the same as those from the authors, while education is given in levels instead of years. 0 implies primary education, 1 secondary or post-secondary education or less and 2 university or higher.

Source: Own elaboration with data set from the INE (Fertility Survey of 2018).

Table A1: Bunching in the number of births around the two cutoffs

Variables	(1) +/- 12 months	(2) +/- 6 months	(3) +/- 3 months
Children born per month around 03/2007	2.326	8.990	13.667
	(5.216)	(5.766)	(5.533)
Children born per month around 01/2017	2.323	1.286	-11.833
	(4.792)	(8.458)	(8.122)

Note: The data comes from the Fertility Survey of 2018. The sample includes all heterosexual fathers and mothers with a child born in a specific time window (indicated in the column headers) around April 2007 and January 2017. The column called +/- 12 months deals with data from April 2006 to March 2008.

Table A2: Balance of covariates around April 2007

Variables	(1) +/- 12 months	(2) +/- 6 months	(3) +/- 3 months
Sex of the kid	-0.055	-0.058	-0.054
	(0.060)	(0.085)	(0.131)
Mother's age	-0.187	-0.327	0.228
	(0.612)	(0.861)	(1.347)
Father's age	1.118	1.582*	1.787
	(0.681)	(0.924)	(1.502)
Mother's education	0.054	-0.155	-0.023
	(0.080)	(0.117)	(0.183)
Father's education	0.077	-0.089	-0.062
	(0.085)	(0.121)	(0.190)
Mother worked before birth	0.003	-0.060	-0.037
	(0.039)	(0.060)	(0.089)
Father worked before birth	0.008	0.006	0.035
	(0.034)	(0.043)	(0.065)
Spanish Mother	0.020	0.028	0.062
	(0.028)	(0.039)	(0.068)
Spanish Father	0.014	0.020	-0.004
	(0.027)	(0.040)	(0.065)
Mother's contract	0.109	0.025	0.015
	(0.068)	(0.093)	(0.153)
Father's contract	-0.063	0.005	0.087
	(0.055)	(0.078)	(0.127)
Mother works in public sector	0.009	-0.036	-0.154
	(0.066)	(0.094)	(0.151)
Father works in public sector	0.052	0.025	-0.101
	(0.061)	(0.090)	(0.135)
Married	0.023	0.083	0.165*
	(0.038)	(0.053)	(0.085)
Mather's age when 1st child born	0.474	0.871	1.589
	(0.669)	(0.917)	(1.419)
Father's age when 1st child born	0.241	1.014	3.521
	(1.473)	(2.393)	(4.270)
Gender norm	0.004	0.012	0.007
	(0.022)	(0.032)	(0.053)
Gender pay gap	0.149	-0.131	-0.541
	(0.238)	(0.334)	(0.503)
N	931	459	236

Note: The data comes from the Fertility Survey of 2018. The sample used includes all heterosexual (non-public worker) fathers and mothers who had a kid with their actual partner in a specific time window of months (indicated in the column headers) around April 2007. The coefficients are from a dummy variable, the main independent variable, which equals 1 if the child was born after the introduction of the reform and therefore eligible for two weeks of paternity leave. The dependent variable is the row header, and each coefficient is computed using a different regression (standard errors in parentheses). All the regressions control for a linear trend in the running variable (the month of birth centered at 0 in April 2007) and enable different trends before and after the reform date. The variable of father public worker has been computed with the complete sample before deleting it, while the rest of the variables are reported without it as it is the main specification used. The number of observations is reported for the specific time window, varying slightly for each variable due to missing observations.

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

Table A3: Balance of covariates around January 2017

Variables	(1) +/- 12 months	(2) +/- 6 months	(3) +/- 3 months
Sex of the kid	0.012	-0.005	-0.002
	(0.064)	(0.092)	(0.138)
Mother's age	0.487	1.033	1.652
	(0.645)	(0.948)	(1.509)
Father's age	0.263	-0.217	1.217
	(0.772)	(1.144)	(1.878)
Mother's education	-0.002	-0.216*	-0.304*
	(0.082)	(0.119)	(0.179)
Father's education	-0.028	-0.109	-0.288
	(0.085)	(0.125)	(0.184)
Mother worked before birth	0.011	0.019	-0.009
	(0.039)	(0.055)	(0.081)
Father worked before birth	0.026	0.066	0.063
	(0.040)	(0.055)	(0.078)
Spanish Mother	-0.072*	-0.066*	-0.044
	(0.030)	(0.036)	(0.049)
Spanish Father	-0.006	-0.000	0.026
	(0.032)	(0.041)	(0.059)
Mother's contract	0.020	0.071	0.108
	(0.074)	(0.103)	(0.163)
Father's contract	0.075	0.083	0.093
	(0.065)	(0.090)	(0.127)
Mother works in public sector	0.029	0.037	0.004
	(0.073)	(0.105)	(0.162)
Father works in public sector	0.072	0.103	0.058
	(0.063)	(0.090)	(0.132)
Number of children	0.060	0.056	0.021
	(0.101)	(0.140)	(0.198)
Married	-0.029	-0.055	0.013

	(0.058)	(0.084)	(0.127)
Mather's age when 1st child born	0.132 (0.776)	0.440 (1.134)	-0.033 (1.649)
Father's age when 1st child born	-0.241 (1.645)	-1.566 (2.634)	-0.222 (5.123)
Gender norm	-0.024 (0.021)	-0.031 (0.030)	-0.091** (0.040)
Gender pay gap	0.021 (0.241)	-0.377 (0.351)	-0.077 (0.538)
N	960	488	247

Note: The data comes from the Fertility Survey of 2018. The sample used includes all heterosexual parents who had a kid with their actual partner in a specific time window of months (indicated in the column headers) around January 2017. The coefficients are from a dummy variable, the main independent variable, which equals 1 if the child was born after the introduction of the reform and therefore eligible for four weeks of paternity leave. The dependent variable is the row header, and each coefficient is computed using a different regression (standard errors in parentheses). All the regressions control for a linear trend in the running variable (the month of birth centered at 0 in January 2017) and enable different trends before and after the reform date.

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

Table A4: Descriptive statistics of childcare time 2007

Variables	Number of observations	Mean	Standard deviation
Panel A: Before April 2007			
Dress	241	0.365	0.482
Buying clothes	232	0.470	0.500
Play	459	0.773	0.419
Take care when sick	476	0.326	0.469
Help with homework	443	0.445	0.497
Choose the food	458	0.325	0.469
Take to school	427	0.494	0.500
Bathe	232	0.470	0.500
Lay down	294	0.605	0.490
Choose extracurricular activities	423	0.697	0.460
Panel B: After April 2007			
Dress	228	0.359	0.481
Buying clothes	437	0.394	0.489
Play	425	0.809	0.393
Take care when sick	441	0.392	0.489
Help with homework	413	0.506	0.501
Choose the food	433	0.358	0.480
Take to school	399	0.511	0.501
Bathe	245	0.473	0.500
Lay down	293	0.594	0.492
Choose extracurricular activities	400	0.730	0.445

Note: The data comes from the Fertility Survey of 2018. The sample includes heterosexual (non-public worker) fathers and mothers who had a kid with their current partner one year before (Panel A) and after April 2007 (Panel B). All the variables are binary, equaling 0 if it is the mother who oversees the activity and 1 if it is both or only the father. The number of observations varies severely for each variable due to the different involvement of parents in different activities as children are to a certain extent self-sufficient

Table A5: Descriptive statistics of childcare time 2017

Variables	Number of observations	Mean	Standard deviation
Panel A: Before January 2017			
Dress	478	0.500	0.500
Buying clothes	486	0.418	0.494
Play	488	0.793	0.406
Take care when sick	459	0.481	0.500
Help with homework	478	0.644	0.479
Choose the food	469	0.345	0.476
Take to school	447	0.600	0.491
Bathe	489	0.630	0.483
Lay down	487	0.597	0.491
Choose extracurricular activities	482	0.749	0.434
Panel B: After January 2017			
Dress	445	0.499	0.501
Buying clothes	452	0.451	0.498
Play	454	0.837	0.370
Take care when sick	443	0.485	0.500
Help with homework	433	0.665	0.472
Choose the food	448	0.344	0.475
Take to school	419	0.611	0.488
Bathe	447	0.638	0.481
Lay down	449	0.626	0.484
Choose extracurricular activities	443	0.765	0.424

Note: The data comes from the Fertility Survey of 2018. The sample includes heterosexual fathers and mothers who had a kid with their current partner one year before (Panel A) and after January 2017 (Panel B). All the variables are binary, equaling 0 if it is the mother who oversees the activity and 1 if it is both or only the father.

Table A6: Outcomes per group 2007

Variables	(1) Egal paygap	(2) Inter paygap	(3) High paygap	(4) Egal GZ	(5) Inter GZ	(6) High GZ	(7) Low gender norm	(8) High gender norm
Mother full time worker	1.371*	0.562	0.652	0.335	0.163	2.033**	0.904*	0.643
	(0.779)	(0.605)	(1.142)	(0.529)	(0.920)	(0.822)	(0.529)	(0.558)
Weeks of mother's leave	-0.957	1.908	-1.690	0.488	-0.077	-0.113	0.197	0.438
	(1.475)	(1.791)	(4.913)	(1.742)	(2.169)	(2.121)	(1.177)	(2.180)
Weeks of father's leave	0.914	1.298	-0.940	-0.786	1.772	1.817	1.360	0.410
	(0.969)	(0.896)	(1.889)	(1.043)	(1.361)	(1.358)	(0.895)	(1.207)
N	307	311	224	373	126	248	464	467
Number of children people who	0.081	0.013	0.378	0.174	0.175	0.392	0.325**	-0.045
had first kid older than 30	(0.176)	(0.183)	(0.240)	(0.138)	(0.266)	(0.249)	(0.141)	(0.154)
N	149	154	120	230	70	90	254	220
Dress	-0.004	0.465	1.886	0.420	0.102	0.464	0.521	0.200
	(0.666)	(0.791)	(1.196)	(0.616)	(1.364)	(0.759)	(0.655)	(0.567)
Buying clothes	0.809	0.824	0.306	0.645	-0.808	0.778	0.657	0.590
	(0.503)	(0.532)	(0.693)	(0.441)	(0.841)	(0.564)	(0.447)	(0.424)
Play	0.882	0.191	1.280	0.717	-0.398	1.284**	0.622	1.083*
	(0.704)	(0.639)	(0.857)	(0.601)	(1.206)	(0.648)	(0.523)	(0.604)
Taking care when sick	0.807	0.804	0.364	0.642	-0.808	0.739	0.630	0.622
	(0.502)	(0.532)	(0.690)	(0.440)	(0.841)	(0.563)	(0.446)	(0.424)
Help with homework	1.204**	0.731	0.160	0.576	0.584	1.061*	0.730*	0.755*
	(0.508)	(0.535)	(0.666)	(0.447)	(0.784)	(0.572)	(0.432)	(0.440)
Food	-0.219	0.619	1.007	0.095	-0.511	0.438	0.046	0.261
	(0.504)	(0.544)	(0.838)	(0.445)	(0.832)	(0.587)	(0.421)	(0.466)
Choose extracurricular activities	0.158	-0.144	0.461	-0.108	0.185	0.406	1.077**	-0.383
	(0.560)	(0.618)	(0.728)	(0.517)	(0.999)	(0.584)	(0.506)	(0.485)
Satisfaction in childcare sharing	1.527**	0.321	1.307	0.810	0.784	1.106	0.930*	0.721
mother	(0.604)	(0.724)	(0.860)	(0.555)	(0.929)	(0.673)	(0.523)	(0.575)
Satisfaction in house chores	0.740	-0.002	1.718*	0.864	0.237	0.066	1.097**	-0.197
sharing mother	(0.736)	(0.777)	(0.939)	(0.620)	(0.980)	(0.788)	(0.554)	(0.686)
Satisfaction with relationship	0.360	0.095	0.684	0.387	0.620	-0.209	0.572*	-0.132
mother	(0.413)	(0.396)	(0.445)	(0.344)	(0.505)	(0.399)	(0.316)	(0.335)
N	253	260	189	305	108	230	399	385
Satisfaction in childcare sharing	-0.813	0.234	0.130	-1.611**	-2.680**	-0.731	-1.336*	0.392
father	(0.908)	(0.872)	(1.333)	(0.782)	(1.118)	(1.667)	(0.739)	(0.704)
Satisfaction in house chores	-1.297	-0.697	0.703	-1.802**	-3.585*	1.687	-1.086	-0.569
sharing father	(0.970)	(0.882)	(1.478)	(0.788)	(1.932)	(1.287)	(0.659)	(0.843)
Satisfaction with relationship	-0.828	-1.335*	0.130	-1.712***	-0.817	0.485	-1.170*	-0.404
father	(0.666)	(0.751)	(1.333)	(0.599)	(1.147)	(0.707)	(0.598)	(0.522)
N	54	51	35	68	18	18	65	82
Percentage of housework done	-2.420	-0.017	-13.998**	-1.607	-6.192	-6.942	-3.387	-3.342
by mother	(4.649)	(4.670)	(6.786)	(4.325)	(8.347)	(5.490)	(4.137)	(4.468)
Percentage of housework done	8.733**	4.681	2.568	1.424	-4.703	9.323*	2.836	3.164
by father	(4.430)	(4.063)	(5.270)	(3.810)	(7.098)	(4.943)	(3.530)	(4.069)

Note: The data comes from the Fertility Survey of 2018. The sample used includes heterosexual (non-public worker) fathers and mothers who had a kid with their actual partner 1 year around April 2007, but varies according to group specification (indicated in the column headers). Columns 1-3 correspond to paygap division, columns 4-6 to GZ grouping and columns 7 and 8 to gendernorm splitting. The coefficients are from a dummy variable, the main independent variable, which equals 1 if the child was born after the introduction of the reform and therefore eligible for two weeks of paternity leave. The dependent variables (those that were found significant or interesting in the sign in the previous results subsections) are the row header, and each coefficient is computed using a different regression (standard errors in parentheses). All the regressions control for a linear trend in the running variable (the month of birth centered at 0 in April 2007) and enable different trends before and after the reform date. Additional controls are the age of the father, being married and the order number of the child in the child table. The number of observations is reported per pools of variables, as the group of people used varies (in order: everyone, everyone if the mother had the first child when she was older than 30, only female respondents, and only male respondents). If the section of the table doesn't have a N, it means that it includes everyone. *** p<0.01, **p<0.05, *p<0.1.

Table A7: Outcomes per group 2017

Variables	(1) Egal paygap	(2) Inter paygap	(3) High paygap	(4) Egal GZ	(5) Inter GZ	(6) High GZ	(7) Low gender norm	(8) High gender norm
Weeks of father's leave	2.761***	0.749**	2.967***	1.692***	2.577***	1.898***	2.021***	1.557**
	(0.803)	(0.373)	(0.869)	(0.459)	(0.959)	(0.432)	(0.429)	(0.632)
N	325	359	178	390	118	270	492	468
Dress	0.708	0.527	-0.287	0.643	0.426	0.668	0.278	0.675
	(0.492)	(0.467)	(0.721)	(0.442)	(0.826)	(0.550)	(0.376)	(0.435)
Play	1.204*	0.626	0.605	0.562	1.684	0.399	1.289**	0.221
	(0.659)	(0.656)	(0.809)	(0.596)	(1.088)	(0.674)	(0.520)	(0.500)
Lay down	0.822	0.191	0.645	0.692	-0.190	0.768	0.468	0.606
	(0.505)	(0.473)	(0.711)	(0.456)	(0.833)	(0.560)	(0.386)	(0.447)
Choose extracurricular	-0.177	-0.939	-0.076	0.265	-1.264	-0.247	-0.071	-0.173
activities	(0.551)	(0.587)	(1.121)	(0.536)	(1.079)	(0.672)	(0.455)	(0.492)
Satisfaction in childcare	0.363	-0.054	-0.520	-0.477	1.513	0.167	0.176	-0.154
sharing mother	(0.532)	(0.593)	(0.806)	(0.530)	(1.073)	(0.670)	(0.440)	(0.568)
Satisfaction in house chores	0.687	-0.225	-0.421	-0.002	0.461	-0.627	0.161	-0.194
sharing mother	(0.609)	(0.611)	(1.021)	(0.571)	(1.284)	(0.722)	(0.470)	(0.653)
Satisfaction with	0.236	0.005	-0.520	-0.325	1.194*	0.453	0.222	0.064
relationship mother	(0.321)	(0.338)	(0.806)	(0.339)	(0.700)	(0.359)	(0.280)	(0.310)
N	264	297	146	302	94	243	426	366
Satisfaction in childcare	0.659	-1.095	-1.772	-0.946	0.470	0.479	-0.408	-0.504
sharing father	(1.247)	(0.821)	(1.752)	(0.871)	(2.649)	(1.504)	(0.942)	(0.847)
Satisfaction in house chores	0.025	-0.992	1.442	-1.173	-0.601	0.791	-0.300	0.313
sharing father	(1.088)	(0.879)	(1.346)	(0.925)	(1.212)	(1.558)	(0.910)	(0.728)
Satisfaction with	-0.460	0.005	0.276	-0.584	1.316	-1.465	0.069	-0.004
relationship father	(0.768)	(0.594)	(0.874)	(0.597)	(0.876)	(1.080)	(0.490)	(0.578)
N	61	62	32	88	24	27	66	102
Percentage of housework	-4.572	5.931	1.128	-4.517	3.486	6.832	0.254	-0.641
done my mother	(4.205)	(4.180)	(6.329)	(3.851)	(8.110)	(5.013)	(3.292)	(4.323)
Percentage of housework	8.168**	-1.436	-7.093	4.729	-8.207	1.951	2.315	0.216
done by father	(3.753)	(3.843)	(5.550)	(3.481)	(7.238)	(4.464)	(2.887)	(3.980)

Note: The data comes from the Fertility Survey of 2018. The sample used includes heterosexual fathers and mothers who had a kid with their actual partner 1 year around January 2017, but varies according to group specification (indicated in the column headers). Columns 1-3 correspond to paygap division, columns 4-6 to GZ grouping and columns 7 and 8 to gendernorm splitting. The coefficients are from a dummy variable, the main independent variable, which equals 1 if the child was born after the introduction of the reform and therefore eligible for four weeks of paternity leave. The dependent variables (those that were found significant or interesting in the sign in the previous results subsections) are the row header, and each coefficient is computed using a different regression (standard errors in parentheses). All the regressions control for a linear trend in the running variable (the month of birth centered at 0 in January 2017) and enable different trends before and after the reform date. Additional controls are the mother's education, the mother's nationality, and the order number of the child in the child table. The number of observations is reported per pools of variables, as the group of people used varies (in order: everyone, only female respondents, and only male respondents). If the section of the table doesn't have a N, it means that it includes everyone. *** p<0.01, **p<0.05, *p<0.1.