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Fathers' Use of Childbirth Leave in Spain. The Effects of the 13-Day Paternity Leave

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Abstract This article investigates, for the case of Spain, to what extent the introduction in March 2007 of a non-transferable 13-day paternity leave has encouraged men to make greater use of childbirth leave. Data were drawn from the Spanish Economically Active Population Survey, covering the period 2005–2009. We use a natural experiment approach, comparing the behavior of wage earners fathers with children of less than 1 year of age before and after the reform and using mothers as control group. After estimating a difference-in-differences logistic regression model we obtain statistical evidence that there is a higher percentage of males on leave in the reference week in the post-reform period (after 2007). The article also analyzes some of the personal and socio-economic determinants of the fathers' use of childbirth leave. Fathers are more likely to be on leave if they have stability in employment, if there are facilities for reconciling work and family life (working in the public sector) and if the partner is employed. The father's age has an interesting U-shaped influence.

Keywords Paternity leave · Natural experiment · Reform of leave system · Fathers' involvement in childcare · Balancing working/family life

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Introduction

Over the past decades, men have been increasingly involved in providing unpaid care (Fursman and Callister 2009). In Spain, the time spent by men daily in domestic work and childcare was on average 1.45 h in 2003, whereas in 2009 it stood at 1.73 h (National Statistics Institute, NSI). This is a phenomenon that occurs slowly, gradually, with different degrees of intensity, and one that occurs more in the most economically and socially advanced countries (in Germany and Sweden, time spent by men on an average day in domestic work and childcare was 2.21 h and 2.29 h, respectively in 2003; see European Communities 2004). In fact this phenomenon is correlated with the fact that women have been entering the labor market and the areas of economic, political, and social power in ever increasing numbers. For example, in Spain, the female activity rate has gone from 43.84 % in 2003 to 51.57 % in 2009 (NSI).

In line with the above, gender equality policies developed by governments, in many cases, have added to their aims that of promoting greater involvement of fathers in childcare and, in general, taking charge of unpaid family work to an increasing extent. One way of moving forward in this objective consists of reforming the parental leave system so that the same possibilities are offered to fathers and mothers to take time off for childbirth. Along these lines a series of countries, such as Iceland (Gíslason 2007; Eydal 2008), Germany (Nitsche 2008; Erler 2009), and Slovenia (Korintus and Stropnik 2009), have over the last decade made a series of reforms in their leave policies, in order to produce a convergence in the rights of childbirth leave for both mothers and fathers, and at the same time promote the use by men of these new rights.

There are several reasons why extending paternity leave would lead to greater fathers' involvement with children. In the first place, the movement toward an egalitarian parental leave system would have an exemplary effect, in the sense that it would eliminate the female bias that was previously in these systems; secondly, it would enable many fathers wishing to take an active part in looking after their young children to do so from the beginning with fewer restraints; and, thirdly, maybe some fathers who initially might have accepted taking paid leave simply to avoid wasting these advantages (2 weeks of paid leave) would end up changing their mind in favor of greater involvement, as a consequence of the new experience they enjoy while looking after their small children. Seward et al. (2006) and Nepomnyaschy and Waldfogel (2007), provide empirical evidence that indicates that fathers who take leave tend to become more involved in childcare.

Advances in the aim of extending paternity leave, at the same time as advances are taking place in other aspects, such as balancing working and family life for mothers/fathers, or the gradual decline of gender stereotypes, may be determining factors in reducing gender inequalities against women that still persist in the labor market (Reich 2010), as well as encouraging more births in Europe (Esping-Andersen 2007; Lappegård 2008). Furthermore, the fact that fathers take full part in the development and care of their children is desirable in itself: it is good for men insofar as it enriches their personal life and even their physical health (Connell 1995; Månsdotter et al. 2007; Malmi 2009), and is good for children since it is very positive for them when their parents spend more time with them (Sarkadi et al. 2007); and is probably good for the parents' stability as a couple (OECD 2007, p. 59).

Spain forms part of those countries in which interest is being manifested in reforming the parental leave system, to give fathers a greater scope. However, these reforms are being introduced slowly and are less wide-ranging than in other countries, such as Nordic ones and Germany (Albert et al. 2008; Castro and Pazos 2008; Ray et al. 2008; Kamerman and Moss 2009; Romero-Balsas et al. 2013). In March 2007, a non-transferable 13-day paternity leave was introduced for the first time, and the government planned to increase its duration to 1 month from January 2011. However, this latter move was deferred as a result of the economic crisis.

One of the aims of this article is to investigate to what extent the introduction of this new 2-week permission has encouraged men to take childbirth leave to a greater extent when they are fathers. Specifically, we wish to find out how far the women–men gap in using leave has been reduced as a result of the introduction of paternity leave. To this end a sample from the Economically Active Population Survey (EAPS) was used. The four quarters of the EAPS were used for the years 2005–2009 both inclusive, by treating them as a cross-section independent data pool in such a way that a sufficiently representative sample was obtained. With that it is possible to ascertain the number of male and female workers off work because of childbirth in the reference week, in each of the quarters analyzed.

From a methodological standpoint the introduction of a fortnight's paternity leave in the first quarter of 2007 also provides a case of a natural experiment. Fathers of children born immediately after the reform are treated differently to fathers of children born immediately before the reform. From a logistic regression model, and using difference-in-differences (DDs) estimation, empirical evidence is given in this article that the introduction of paternity leave in Spain has reduced the woman–man gap in the use of leave periods, even though this gap is still very wide.

Moreover, it is worthy of note that male workers who make use of childbirth leave are for the most part those with the most stable jobs in the firm (they are on permanent contract); those who work in organizations providing the most facilities for reconciling work and family life (public sector); those who live in regions where public policy encourages particularly the use of leaves; and those who have fewer economic restrictions to do it (they have employed spouses). This suggests that gender ideology may not be something that is fixed and does not fully determine the individual decision to take leave. This decision also may depend on the constraints and incentives system in which the father operates.

For the Spanish case, this study would be the first that analyzes the determinants of men's use of childbirth leave from a sample deriving from a macro-official survey (EAPS). Lapuerta et al. (2011) also analyzed these determinants for the case of Spain, but they studied the specific case of Spanish unpaid parental leave. As we will show later, our results can be compared with those existing for other countries (e.g., Eriksson 2005, for Sweden; Gíslason 2007, for Iceland; Nepomnyaschy and Waldfogel 2007, for The United States; and Reich 2010, for Germany).

Reform of the System of Parental Leave in Spain

Since the 1980s, several European countries have established different types of paternity leave or have reserved a part of the parental leave for the parent who does not take most of the leave, which is the father in most cases (Kamerman and Moss 2009). The case of Spain corresponds to the former. In March, 2007, paternity leave was introduced although of modest length of time (Wall and Escobedo 2009; Escobedo and Meil 2012).

Paternity leave is for 13 days, uninterrupted (for large families or households with a disabled person the leave is extended by 5 days; two more days are added in the case of multiple births, or multiple adoption or fostering). It can be taken full-time or part-time. It can be taken from the day of the birth (or adoption or fostering of a minor), till immediately after the end of the maternity leave. Paternity benefit amounts to 100 % of earnings. If the paid childbirth leave (2 days paid by the firm), is taken into account, Spanish wage earners fathers can be considered to be entitled to 15 natural days (approximately a fortnight) of paid paternity leave.[1] The 13-day paternity leave is considered as a transition toward a month's paternity leave, which was foreseen as coming into force in January 2011. However, the Spanish government has postponed its entry because of the economic crisis.

Moreover, there is maternity leave, which lasts for 16 continuous weeks (when there is a multiple birth two more weeks are added for each baby, after the second). This period can be taken on a full or part-time basis. The mother can transfer part of her leave to the father, except for the first 6 weeks immediately following the birth, which she is obliged to take. The father takes these weeks transferred, either successively or simultaneously with the mother's rest. Maternity benefit represents 100 % of earnings.

Furthermore, each parent is entitled to take unpaid parental leave for a maximum 3-year period after childbirth (leave is an individual right). There is no payment, but some regional governments offer low flat-rate benefits (Lapuerta 2013). For example, the Basque Country offers €200 per month for mothers and €250 per month for fathers in 2009; La Rioja, Navarra, Castilla-La Mancha, Castilla y León, and Murcia offer similar benefits. There is also unpaid leave for a maximum period of 2 years for looking after a dependent family member.[2]

[1] Wage-earning men (those belonging to the sample for this article) have 15 days paid paternity leave (13 days paternity leave plus 2 days paid childbirth leave). Self-employed men have only the 13 days paternity leave. To obtain the 13-day paternity leave it is necessary to have made a minimum of 180 days contributions to the Social Security (within the 7 years immediately preceding the date at which the leave commences). These eligibility conditions are not very restrictive. Thus, it is to be expected that the vast majority of wage-earning men included in the sample for this work are eligible for paternity leave.

[2] There are two additional employment-related measures (Escobedo and Meil 2012): "permiso de lactancia" ("breastfeeding leave"), according to which during the first 9 months after the child's birth employed mothers (wage earners) are entitled to 1 h of absence during the working day without loss of earnings, which is paid by the employers. If both parents are working, the mother can transfer this right to the father. All employed mothers can consolidate this reduction in working time as full-time leave, thus in practice extending their maternity leave by between 2 and 4 weeks. And "reducción de jornada por cuidado de un hijo", according to which a working parent (wage earner) can reduce his/her working day by between an eighth and half of its normal duration to care for a child until the eighth year or to look after a disabled child.

Table 1 Take-up of leave by sex, in Spain

	2005	2006	2007	2008	2009
Births	466,371	482,957	492,527	519,779	494,997
Maternity leave					
Females (% births)	296,115 (63.5)	317,318 (65.7)	326,438 (66.3)	353,585 (68.0)	334,786 (67.6)
Males (% births)	5,269 (1.1)	5,282 (1.1)	5,204 (1.1)	5,575 (1.1)	5,726 (1.2)
% Males/females	1.8	1.7	1.6	1.6	1.7
Paternity leave					
Males (% births)	–	–	173,161 (35.2)	279,756 (53.8)	273,024 (55.2)
Unpaid parental leave					
Females (% births)	27,457 (5.9)	30,052 (6.2)	33,335 (6.8)	36,300 (7.0)	32,549 (6.6)
Males (% births)	946 (0.2)	1,223 (0.3)	1,481 (0.3)	1,471 (0.3)	1,393 (0.3)
% Males/females	3.4	4.1	4.4	4.1	4.3
Unpaid leave caring of family member					
Females	2,813	3,336	4,575	4,784	4,534
Males	519	594	846	880	768
% Males/females	18.5	17.8	18.5	18.4	16.9

Note Data on leaves come from the Spanish Social Security (Ministerio de Empleo y Seguridad Social, Anuario de Estadísticas). Data on births come from the National Statistical Institute (Cifras de Población y Censos Demográficos)

Table 1 presents the number of people who started a statutory maternity or paternity benefit as well as those who started an unpaid parental leave (these are the only data available in the official statistics. There are no official data on the duration of leave). The table also shows the number of births that took place in Spain in each year, in order to calculate the rate of use of leaves with respect to total births (Escobedo and Meil 2012). The period considered is 2005–2009, which is the one used in the quantitative part of this article and includes the 2 years prior to and following the introduction of paternity leave in 2007.

As can be seen, only a small number of males made use of maternity leave compared to the number of women taking advantage of it. Throughout the period under consideration only about 1.71 % of maternity leave was transferred to the man.

It must be highlighted that in Spain there is fairly low use of unpaid parental leave for looking after children (for example, in 2009, 334,786 women used maternity leave, whereas only 32,549 made use of unpaid parental leave). Moreover, the duration of unpaid parental leave in most cases does not exceed the first year after the birth (Lapuerta et al. 2011). And once again, men are very much in the minority of users of unpaid parental leave. Between 2005 and 2009 the ratio males/females who started an unpaid parental leave was slightly over 4 %.

Unequal use of paid and unpaid leaves is somewhat reduced when unpaid leave for caring for family members is considered. In this case, the ratio males/females is approximately 18 %.

Finally, the introduction of paternity leave has meant an important change in the extent to which men make use childbirth leaves. In 2007, the first year in which paternity leave was in force, the number of men using it was 173,161, covering 35.2 % of the 492,527 births in this year (in the months of January and February this leave was not available); in 2008, this figure rose to 279,756 (53.8 % of the births), and in 2009 reached 273,024 (55.2 %). Judging from these data, the introduction of paternity leave has been a success, given that it is being used by a majority of men (of those entitled to it).

In this article, as will be explained below, we use microdata from the EAPS. These data make it possible to know how many women and men say they are on childbirth leave in the reference week. In terms of the previous discussion on the Spanish leave system, that means that the fathers that are on leave in the reference week, technically could be on it by using maternity or unpaid parental leave, before 2007; or by using paternity, maternity, or unpaid parental leave, after the first quarter of 2007.

Moreover, in the case of Spain, the study period coincides with the implementation of a rather ambitious series of public policies on gender equality (Valiente 2008). The most important measure of this kind was the introduction of the "Law for Equal Opportunity between Women and Men", in March 2007. In fact, the 13-day paternity leave was one of the points of this law. This law also introduced a new baby bonus for each birth or adoption, that was in force between July 2007 and December 2010. This bonus was received by the mother and its purpose was to compensate for some of the economic cost associated with the birth or adoption. Since it was not linked to the fact of being in the labor force or using the leave system (it was a bonus for every birth or adoption), in this article it makes sense to assume that it had no influence in the way parents behaved toward leave.

Finally, from 2008, the effect of the economic crisis may have negatively influenced the willingness of fathers to take leave. Reduction in staff and increased unemployment has been particularly severe in Spain, and this may have led many male (and female) workers to feel that their jobs were at risk. As a result, some of them may have become more reluctant to take leave that they were entitled to. In order to control for this possibility, the logistic regression analysis will consider the unemployment rate (in the province of residence) as a proxy of the deterioration of the economic environment of the surveyed person.

Review of Literature on Evaluating the Changes in Leave Policy and Working Hypothesis

A series of studies can be highlighted in which an evaluation is made of the introduction of a substantial reform in the leave system, normally including some extra incentive to persuade men to use this leave. They all attempt, among other things, to analyze how much the introduction of the reform has led to an increase in the extent to which men use these periods of leave.

The Icelandic Act on Maternity/Paternity and Parental Leave underwent significant changes in the year 2000 (Eydal 2008, 2009; Einarsdóttir and Pétursdóttir 2009). The leave was extended from 6 to 9 months, and in addition was distributed so that fathers were given 3 months' leave, mothers 3 months and the parents were given 3 months to share as they wished. This means that the reform introduced in Iceland equalized legal rights to paid parental leave. Gíslason (2007) analyzes the effect of the introduction of this reform in Iceland. He uses several kinds of data (from general official statistics to surveys by private companies). The proportion of fathers utilizing their leave in whole or in part was already very high in the first year (82.4 %) and kept growing up to a rate of around 90 % in 2004. In the same way, the average number of days fathers were taking was growing and was clearly linked to the non-transferable right of fathers (fathers used on average 96 days, while mothers used an average of 182 days in 2004). Gislason shows that, compared to the other Scandinavian countries, the proportion of fathers using their right to paternity leave is the highest in Iceland.

Sweden was one of the first countries to earmark part of the parental leave for fathers. One "daddy month" was introduced in 1995 (Chronholm 2009). In 2002, the number of months reserved for fathers in the Swedish parental leave system increased from one to two. At the same time there was an increase in the total time of parental leave from 12 to 13 months. Eriksson (2005) analyzes how this policy change (the introduction of the second daddy month) affected the fathers. From a methodological point of view, the daddy month reform allows the use of a natural experiment approach, because parents of children who are born after 1 January 2002 are affected by the reform, but not parents of children born before. He uses data from the registers of the Swedish National Social Insurance Board and from the Swedish Standard-of-Living Survey. Eriksson concludes that both fathers and mothers increased their use of parental leave after the reform. In particular, the mean increase in the use of parental leave by fathers was the result of a decrease in the number of fathers using 1 month or less and an increase in the number of fathers using about 2 months of parental leave.

In 2006, Germany passed a new parental leave policy which then came into effect in January 2007. This new German parental leave scheme is similar to the Swedish model of family policy (Erler 2009). It provides for parents, after childbirth, taking paid leave of up to 14 months in order to care for their newborn child. Although either of the two parents may take the first 12 months of the leave, the 13th and 14th month are reserved for the parent who did not take the first 12 months of the leave. These are the so-called “daddy months”, and can therefore only be taken when the leave is shared between both parents. Reich (2010) shows that this reform led to a strong increase in the share of leave taken by fathers (in 2009 fathers used about 18 % of the time corresponding to allowances for parental leave). She also tries to identify the socioeconomic and workplace-related determinants of the fathers’ use of parental leave after the introduction of this new parental leave system, by using the 2007 German Microcensus as a database, and estimating three logistic models. She concludes that the fathers’ use of parental leave is higher among fathers with a secure job (those with a permanent contract), and among those working in the public sector. There is also strong dependence on the female partner’s net wage and employment status.

Han et al. (2009) analyze the case of the US. They describe trends in leave-taking after birth of a child and analyze the extent to which these behaviors are influenced by parental leave policies. These parental policies include the introduction of the Family and Medical Leave Act in August 1993, which provides up to 12 weeks of unpaid leave for specified reasons, including the birth or assumption of care of a new child. They also consider the case of several states that enacted parental leave laws separately from the federal legislation. They use data from the June Current Population Survey (CPS) Fertility Supplements, merged with other months of the CPS, and cover the period 1987–1994. Then they estimate a series of econometric models of DDs. Their main findings suggest increased leave-taking by both mothers and fathers. The magnitudes of the changes are small in absolute terms, but large relative to the baseline for men.

In this work, following the line of the articles just cited, we wish, first, to test the following hypothesis:

Hypothesis 1 Fathers respond to parental leave policies.

The introduction of paternity leave of 13 days in Spain has increased the probability of wage earner fathers with children of less than 1 year of age being on leave for childbirth in the reference week.

The present study also explores some of the determinants of fathers’ use of childbirth leave.

Partly by following Lapuerta et al. (2011), five interrelated factors can be distinguished when studying the determinants of fathers’ use of childbirth leave (we develop these factors in the subsequent paragraphs). These are as follows: *personal circumstances that facilitate the reconciliation* of working and family life (for example, working in the public sector, where it is easier to take leave). The *gender ideology* (the more egalitarian the gender attitudes the greater the male's willingness to take leave). The *household or couple characteristics* that, according to bargaining models (Lundberg and Pollak 1996), influence the father's willingness to take leave. The *opportunity cost* (a higher opportunity cost results in a lower inclination to take leave). And *work/family reconciliation policies* (if there is an extensive parental leave system, a high level of social and childcare coverage, and organizations are aware about the importance of being family-friendly companies, there will be more facilities for men to take leave).

Each of these five factors to a greater or lesser extent, and simultaneously, helps to explain why a certain type of father is more or less willing to take childbirth leave. For example, the probability that a highly skilled professional father will take leave will be influenced negatively by the high opportunity cost that he faces (professional opportunities that can be lost during the period of leave). However, this effect may be counteracted by the fact of being young (30 years) and having a high educational level, as this increases the probability of having more egalitarian gender attitudes; by the fact of working in the public sector, which is a family-friendly organization; by the fact of having a wife who is also a skilled professional; and because now the government is developing a campaign to encourage fathers to use the parental leaves.

The literature on the determinants of the fathers' use of leave proposes a set of explanatory variables, most of which come within the framework of one of the above five factors.

Several studies show that men employed in the administration take childbirth leave more often than those employed in the private sector (Brandth and Kvande 2002; O'Brien and Shemilt 2003; Lammi-Taskula 2008; Reich 2010). That is to say, it is a case of *personal circumstances that facilitates the reconciliation.* The reason for this may be found in the greater stability and better working conditions in the public sector, where it may be easier for fathers to exercise their statutory leave rights (Lammi-Taskula 2008). A similar interpretation is that the public sector is a family-friendly organization that provides organizational work–family initiatives to help employees (women and men) to balance work and family life (Poelmans 2008), and this favorable environment also encourages men to take childbirth leave.

Having an indefinite contract is another case of *personal circumstances that facilitate the reconciliation.* Having an indefinite contract can be used as an indicator of stability and security in the job, and it is to be expected that it will have a positive effect on the use of leave by men (and women). Reich (2010) and Geisler and Kreyenfeld (2009) provide evidence in this sense for the case of Germany. Reich (2010) also points out that not only the father's job security but also the mother's is vital for the leave taking of fathers.

Highly educated fathers often have better jobs with enhanced leave entitlements (*personal circumstance that facilitates the reconciliation*), and, therefore, tend to be more willing to ask for leave (Whitehouse et al. 2007). But better educated fathers also are hypothesized (Yeung et al. 2001) to be more knowledgeable about children's developmental need for positive paternal involvement. Moreover, the higher the level of education, the more likely the father is to hold more egalitarian gender attitudes and modern views on childcare (Smeaton 2006). Expressed in another way, if values and beliefs are important for men's parenting activities, and highly educated individuals are forerunners in terms of new values and ideas (Geisler and Kreyenfeld 2009), highly educated men should be more likely to take leave than their less educated counterparts (*gender ideology factor*). There exists empirical evidence on the positive relationship between fathers' educational level and their propensity to take leave for the cases of Sweden (Sundström and Duvander 2002), Norway (Brandth and Kvande 2002), Australia (Whitehouse et al. 2007), United Kingdom (Smeaton 2006), and Germany (Geisler and Kreyenfeld 2009).

The age of the male may influence his willingness to take childbirth leave. According to Alberdi and Escario (2007), changing gender attitudes in favor of the father's involvement may be more important among young men than mature ones (*gender ideology*). But, at the same time, the youngest men may be least likely to be in an employment situation providing secure access to leave arrangements (Nepomnyaschy and Waldfogel 2007; Whitehouse et al. 2007), and this situation could lead them to a lesser extent to take childbirth leave. This suggests the possibility of a non-linear relationship, with the willingness of men to take leave initially (for the youngest) increasing with age, and then, for most men, potentially declining with age.

Some studies have found that the take-up rates are different for some minority groups in the country. For example, Seward et al. (2006), carried out a study of fathers taking childbirth leave in the US and, among other things, concluded that being a Hispanic father was negatively correlated with taking leave. In the quantitative part, which will be developed below, a dummy variable will be introduced to include the fact of being an economic immigrant. In this population group there may be lower take-up rates because among the males in the group there is a higher incidence of job insecurity along with lower human capital levels, but it may also be that among the members of this group more traditional gender attitudes are prevalent (*gender ideology*).

The willingness to take leave may be more prevalent among males whose partners are working, compared to those whose partner does not work (*household or couple characteristics*). According to bargaining models (Manser and Brown 1980; McElroy and Horney 1981; Sen 1990; Lundberg and Pollak 1996), the power relations between spouses affect the household division of labor. The relative power of partners derives from their command over resources, which are often measured by the earnings of the husband and the wife. Implicit in these models is that household labor is undesirable, and thus performed by those with less power. Therefore, wives can use their earnings to “buy” increased participation by their husbands in household tasks (Yeung et al. 2001). In the present study, there are no data for the earnings but there are data as to whether the partner works or not. If the wife has a job (she is employed in the labor market), the male is not the only one obtaining earnings (they are a dual earner family), so he will have less power in negotiations about the distribution of the housework (including caring for children), and will tend to do more of it. But, on the other hand, the assumption that housework and childcare belong to the same category can be questioned. Household labor is usually regarded as an unpleasant duty, whereas investing time in parenting is considered a more rewarding task (Geisler and Kreyenfeld 2009), and this fact may accentuate the idea that in dual earner families male's willingness to take leave is higher.

Willingness to take leave could be less frequent in the case of managers and high-income males than among the rest of the male workforce (Lapuerta et al. 2011). The time spent in caring for children has an *opportunity cost* (the cost in terms of the best alternative that must be foregone in order to pursue a certain action). The time parents invest in caring for children carries an opportunity cost of both the earnings foregone and the human capital accumulation sacrificed (Mincer and Polachek 1974; Becker 1991). There are trade-offs between investing in children and in themselves. The opportunity cost of taking leave for childbirth (when professional opportunities can be lost) will be considerably greater among workers occupying management or high responsibility posts than among the rest of the workers. Consequently, managers and high-income fathers are expected to spend less time with children (Yeung et al. 2001; Herrarte et al. 2012).

Finally, in Spain there is territorial diversity, expressed in social behavior and political decentralization, affecting many fields relevant to gender and leave policies (Wall and Escobedo 2009). Those regional differences in the economic policies applied may impinge in a differentiating way upon the extent to which men use leave for childbirth. In Spain, in five regions (Basque Country, La Rioja, Navarra, Castilla-La Mancha, Castilla-León, and Murcia) there are some extra benefits to fathers (or mothers) who use the unpaid parental leave (Lapuerta et al. 2011). Also in some of these regions (such as the Basque Country) there has been a particular emphasis on promoting the involvement of fathers in caring for small children (Emakunde 2007; Bergara et al. 2008). Later on, in the quantitative part of the article, a dummy variable will be added to include these regions, to compare their results with those of the rest of the regions.

After consideration of these determinants of fathers' use of parental leave, we propose five additional hypotheses:

Hypothesis 2 Fathers' use of childbirth leave is greater if they have personal circumstances that facilitate the reconciliation.

Fathers (with children of less than 1 year) working in the public sector and those with indefinite contracts have more facilities to take leave. Therefore, a higher probability of being on leave in the reference week can be expected among them.

Hypothesis 3 There is a non-linear relationship between the age of the father and his use of childbirth leave.

Fathers' use of leave initially (for the youngest) would increase with age, and then, for most men, would potentially decline with age.

Hypothesis 4 There is a positive relationship between the level of education of the fathers and their use of childbirth leave.

This would reflect (among other things) that gender attitudes and knowledge about children's developmental need for positive paternal involvement are generally more advanced among fathers with a high education level.

Hypothesis 5 Fathers' use of childbirth leave is lower among economic immigrants.

Among this group of population there may be more traditional gender attitudes that would affect negatively the male take-up rate. Also the job insecurity would affect negatively the (male and female) take-up rate (in our model we will control for some variables related to job insecurity, such us "temporary contract", but not completely; for example, we do not include the fact or having an informal job).

Hypothesis 6 Fathers' use of childbirth leave is greater in dual-earner families.

According to couples' bargaining approach, the probability that fathers with children of less than 1 year of age are on leave in the reference week is higher when they have working partners.

Hypothesis 7 The opportunity cost has a negative effect on fathers' use of childbirth leave.

The probability that fathers with children under 1 year of age are on leave in the reference week is lower among managers, since they usually have a high opportunity cost (especially in the form of high professional opportunities that can be forgone during the period of leave).

Hypothesis 8 Regional differences in leave policies may generate differing take-up rates.

In the five regions (Basque Country, La Rioja, Navarra, Castilla-La Mancha, Castilla-León, and Murcia), where the leave system is slightly more extensive, we expect to find higher take-up rates.

Data and Methodology

In Spain there is no specific statistical project devoted to measuring how people use leave arrangements as well as the characteristics of the people taking leave. The only official survey which indirectly registers the cases of childbirth leave is the EAPS. The EAPS is a continuous quarterly survey that targets households, and its main objective is to obtain data on the labor force, and on the people outside the labor market. The theoretical sample varies from 65,000 households per quarter to around 60,000 actually interviewed households, which implies approximately 180,000 people. All definitions and criteria used are in line with those established by international organizations dealing with labor force surveys and, in particular, they are in accordance with the norms of the European Statistical Office (EUROSTAT).

The reference period for the results of the survey is the quarter. The EAPS is a continuous survey that is made every 3 months, with interviews undertaken during the 12 weeks of each quarter. But the reference period for the information collected is the *reference week* (from Monday to Sunday) prior to the date the interview takes place.

On the basis of two questions asked in the questionnaire it is possible to find out whether in the reference week the worker was (totally or partially) on paid or unpaid leave for the birth of his/her child. In this article, the person interviewed is regarded as totally or partially off work because of the birth of his/her child if the answer to the question "reasons why you did not work in the reference week, when you had a job" is "leave due to the birth of a child" or "unpaid leave due to the birth of a child". The same occurs if the answer to the question reasons why you worked fewer hours than usual" is that it was because of "leave due to the birth of a child".

This information is limited, since it is not possible to discover how long these cases of leave lasted. However, knowing how many male and female workers were off work because of childbirth in the reference week does make it possible to quantify how great was the women–men gap in the use of leave, and how this had been modified following the introduction of paternity leave in 2007.[3]

To have at our disposal a large enough sample of Spanish parents, initially 20 EAPS quarters were selected, covering the years 2005–2009 (there are 2 years before and 2 years after the 2007 reform). These are treated as a pool of independent cross-section data, that is, we treat them as though each sample for each quarter corresponds to a random sample of the working population independent from the previous one (Wooldridge 2002).

From this basis the subsample of wage earners is taken, with ages ranging from 16 to 64, and with children under 1 year. Only wage earners were chosen (self-employed workers were omitted) in order to have a more homogeneous sample of employed workers. Likewise, wage earners with children under 1 year were chosen because in the case of Spain, given the short period of maternity and paternity leave and given the limited take-up of unpaid parental leave, most of the childbirth leave is taken in the first year after the birth (Escobedo and Navarro 2007; Lapuerta et al. 2011).

In Table 2, sample data from 2005 to 2009 are presented disaggregated by quarters. The total sample is of 31,449 wage earners with children under 1 year (12,769 women and 18,680 men). Among them, in the reference week 4,495 women and 462 men were on leave. That means that in the reference week 35.20 % of women with children under one were off work, and only 2.47 % of the men with children under one were in this situation between 2005 and 2009.

Empirical Strategy: The difference-in-differences (DDs) Estimator

We want to test the hypothesis that "the introduction of paternity leave of 13 days in Spain (the treatment) has increased the probability that fathers with children under one year are on leave for childbirth in the reference week" (Hypothesis 1).

To test this hypothesis, it is necessary to perform an analysis of causality (Heckman 2000). We need to use some method of causal inference that allows us not only to study the correlation between policy implementation and the frequency

[3] Note, therefore, that we will focus on the number of fathers and mothers who were on leave in the reference week during a given period (the first year of life of the child) and not on the average durations of these leaves.

Table 2 Sample used and percentage of it on childbirth leave in the reference week

	Both sexes		Females		Males	
	N	Percent on leave	*N*	Percent on leave	*N*	Percent on leave
2005—1°Q	1,436	13.79	545	34.31	891	1.23
2005—2°Q	1,444	13.85	535	34.77	909	1.54
2005—3°Q	1,408	13.28	525	34.29	883	0.79
2005—4°Q	1,443	15.80	559	37.57	884	2.04
2006—1°Q	1,429	13.16	552	32.43	877	1.03
2006—2°Q	1,504	15.89	602	38.04	902	1.11
2006—3°Q	1,551	15.41	614	35.99	937	1.92
2006—4°Q	1,540	13.44	602	32.06	938	1.49
2007—1°Q	1,578	14.45	625	33.92	953	1.68
2007—2°Q	1,689	17.64	676	39.94	1,013	2.76
2007—3°Q	1,628	15.17	652	33.44	976	2.97
2007—4°Q	1,651	15.87	670	34.78	981	2.96
2008—1°Q	1,664	16.53	673	34.18	991	4.54
2008—2°Q	1,666	16.57	696	35.63	970	2.89
2008—3°Q	1,705	17.54	722	36.98	983	3.26
2008—4°Q	1,623	16.64	700	34.29	923	3.25
2009—1°Q	1,585	16.66	696	33.19	889	3.71
2009—2°Q	1,580	16.52	678	33.92	902	3.44
2009—3°Q	1,652	17.86	718	36.77	934	3.32
2009—4°Q	1,673	17.69	729	36.63	944	3.07
Total	31,449	15.76	12,769	35.20	18,680	2.47

Source EAPS

N total sample of wage earners in each quarter with ages ranging from 16 to 64 and with children under 1 year of age

with which fathers use childbirth leave, but also the direction of causality: Is the reform what makes more fathers to take leave and not other factors. In order to evaluate the effectiveness of the introduction of paternity leave we may use an experimental approach, trying to measure how the propensity to take the paternity leave changes for each individual in the sample as a consequence of the reform (Heckman and Robb 1985). Unfortunately, we do not have experimental data. We only have observational data, that is, we observe only if a father (or a mother) is on leave at any given time. All we can do is making comparisons between groups of individuals affected by the policy change and other groups not affected. These methods of causal inference based on use of observational data for the evaluation of public policies are included in the approach of natural experiments (or quasi-experiments; Campbell 1969; LaLonde 1986; Meyer 1995; Imbens and Wooldridge 2009).

In our area, a natural experiment occurs when as a result of a change in policy or law there is also a change in the rules of the game in which individuals operate. To test whether it is the policy change that leads individuals to behave differently and not the simple passage of time or other external factors, there must be some control group not affected by the policy change with which to compare the results of the treatment group affected by the reform.

Unlike a controlled experiment, where homogeneous groups of treatment and control are chosen randomly by the researchers, in natural experiments the groups arise from the change in the specific policy itself. The validity of the comparison between the two groups requires compliance with the assumption of independence between the potential outcomes of the policy and the decision of belonging or not to the treatment group. If some individuals could determine whether they belonged to the treatment group (presumably when they expect a net benefit of treatment), there would be a problem of endogeneity or self-selection in the treatment. When this endogeneity problem is not considered and solved, the estimated effect of the change in the policy on outcome would be biased and the empirical causal inference would not be valid (Heckman and Smith 1995; Besley and Case 2000).

The introduction of a fortnight's paternity leave in the first quarter of 2007 provided a case of a natural experiment. Fathers whose children were born immediately after the reform were treated differently to fathers whose children were born immediately before the reform. There is also a clear candidate—mothers—to become the control group not affected by the policy, because reform in the leave system in March 2007 only affected paternity leave, while the legislation on maternity leave was not altered in this period.

In the present case, we think there is no problem of self-selection. This endogeneity problem could only arise from those very rare cases where parents had decided (more than 9 months before March 2007) to delay the moment of conception in order to give birth after the entry into force of the new paternity leave. However, the entry into force of paternity leave was announced at almost the same time as the approval and coming into force of the law (the law was passed and entered into force simultaneously in March 2007). In any case, we have removed from the sample all the year 2007 to avoid totally the possibility of endogeneity, but also to be sure that the fathers of our sample belong clearly to the group without paternity leaves (years 2005 and 2006), or to the group of fathers with the possibility to take the paternity leaves (years 2008 and 2009).[4]

Once the problem of self-selection has been solved, we will use the DDs estimators, that constitute a common empirical methods used in applied economics for the program evaluation using natural experiment approach (Meyer 1995; Angrist and Krueger 2000). The DDs estimator compares outcomes between a control group

[4] In this sense, the exclusion of the whole of 2007 also reflects the fact that it is possible that some parents with a child under 1 year of age over the sample of the last three quarters of 2007 that we assign to the post-treatment group, had their child before March 2007 and, therefore, that those parents actually belong to the group of pre-treatment with no right to use paternity leave of 13 days. Given the difficulty of clearly identifying the group of parents with children less than 1-year-old who belong to the pre-treatment and those in the post-treatment group during 2007, we decided to completely remove them from the sample (although perhaps the elimination of only the first two quarters of 2007 would have been sufficient we have decided to eliminate all four 2007 quarters).

(mothers), not affected by the policy change, and a treatment group (fathers), that after a certain time is affected by the legislative change.

Considering that both the treatment group (fathers) and control group (mothers) are under the influence of any potential factor unrelated to the legal reform in paternity leave but affecting the propensity of taking parental leave, the DD estimator corrects the simple difference in the outcome before and after within the treatment group (fathers) by subtracting the simple difference in outcome before and after within the control group (mothers) to isolate the effect of policy change.[5] In this way the sample of individuals analyzed may be broken down into four groups: fathers before the reform, mothers before the reform, fathers after the reform, and mothers after the reform.

A simple graphical analysis can initially help us to evaluate the introduction of paternity leaves in a first stage using the DD methodology. We define the take-up rate as a percentage of men (women) taking leave in the reference week compared to the total number of men (women) in that week. In Fig. 1, the evolution of the male and female take-up rates is shown. The 20 quarters comprising the data pool used in this work appear in the figure grouped in 5 years. If a comparison is made before and after the introduction of paternity leave (2007), an increase or leap can be observed in the take-up rate for men. The rate at which leave was used by the men for the period prior to the measure was 1.4 %, whereas the subsequent rate of use was 3.4 %, while the take-up rate for women was maintained roughly constant around 35.2 % in the same period. This appears to suggest that the introduction of paternity leave has had the expected positive effect.

However, the observed change in the male take-up rate could have been caused by any other factor unrelated to the policy change, and for this reason we have to complete this preliminary analysis in order to conclude about causality effects, in particular we have to consider the validity of the parallel trends assumption as we discuss later.

For the moment and in order to quantify the differences observed graphically, the DDs estimator will be implemented in a regression analysis framework. We are interested in evaluating the effect of the introduction of the paternity leave on take-up by the men, as recorded by the variable “leave”, which is a dichotomous variable coded “1” if the person was on childbirth leave (paid or unpaid) in the reference week and “0” in the opposite case.

The DD regression analysis has at least three independent variables. Firstly, is a binary variable “treatment” reflecting the policy change to be evaluated, in our case, the introduction of paternity leave of 13 days in Spain in 2007.[6] This “treatment” is a dichotomous variable that identifies the two periods, the pre-reform (treatment = 0, from the first quarter of 2005 to the fourth quarter of 2006), and

[5] If we limit ourselves to compare only the take-up of fathers before and after the introduction of the paternity leave (without considering mothers), we could be erroneously attributing to that policy any change in the propensity of fathers to use parental leaves, although the origin of this change could be in any other factors unrelated to the reform in paternity leave but which occurred during the same period of time.

[6] “Treatment” refers in our case to the fact that men did not have access to paternity leave till the first quarter of 2007, but they did after that date.

Fig. 1 Percentage of males on leave over total male wage earners and percentage of females on leave over total female wage earners, with ages range from 16 to 64 and with children under 1 year

post-reform (treatment = 1, from the first quarter of 2008 to the fourth quarter of 2009). The second independent variable is the binary variable "male" that identifies fathers (male = 1, the treatment group) and mothers (male = 0, the control group). The third independent variable is the interaction of treatment and male (treatment × male), another binary variable identifying fathers after March 2007 (after the introduction of paternity leave). The coefficient estimated on this male × treatment variable will be isolating the treatment effect, that is, the effect of treatment on the treated, the effect of the introduction of paternity leave on propensity of fathers to take childbirth leaves.

In this case, the DDs treatment effect estimator will be obtained in a logistic regression model to estimate how treatment, male, and their interaction treatment × male help to explain the probability of an individual will take childbirth leave (Wooldridge 2002). Thus, the equation to be estimated to evaluate the policy would be:

$$P(\text{leave} = 1) = \Phi[\beta_0 + \beta_1\text{treatment} + \beta_2\text{male} + \beta_3\text{treatment} \times \text{male}], \quad (1)$$

or alternatively

$$\text{Log}\left[\frac{P(\text{leave} = 1)}{1 - P(\text{leave} = 1)}\right] = \beta_0 + \beta_1\text{treatment} + \beta_2\text{male} + \beta_3\text{treatment} \times \text{male} + u,$$

where u is a purely random variable and $\Phi[\cdot]$ is the logistic probability distribution function. The parameter we use to test whether "the introduction of paternity leave has increased the probability that fathers with children of less than one year of age are on childbirth leave in the reference week" is β_3, associated with the interaction term, "treatment × male". The parameter β_1 associated with the dummy "treatment" captures aggregate factors that affect the probability of "leave = 1" over time in the two periods (treatment = 0 and 1) in the same way for both groups. The presence of "male" by itself captures possible differences between the treatment (fathers) and control (mothers) groups before the policy change occurs.

DDs Treatment Effect Estimator: Preliminary Results

Table 3 shows the estimated logistic regression (1). The results show that men generally have a lower propensity to be on leave than women, and that the treatment variable is only significant when interacting with male, that is, the treatment has only significant effects on men, as we expected.[7] More specifically, the probability that a man, wage earner with a child under 1 year of age, was on childbirth leave increased by 142 % after the legislative reform, growing from 1.4 to 3.4 %. We can conclude then that these first results seem to support the effectiveness of the introduction of paternity leave in 2007, and are not rejecting Hypothesis 1.

Before accepting this result, however, we must ensure that our DDs estimate does not present certain problems, such as "parallel trends", since these problem could be biasing our estimates and consequently invalidating our conclusions about the effectiveness of the introduction of paternity leave.

Results Regarding the Determinants of Fathers' Use of Childbirth Leave: The DDs Estimator Controlling by Observables

One of the most common problems with DDs estimates is the failure of the parallel trend assumption causing DDs estimates in (1) to be biased.[8] In general terms, the parallel trend assumption establishes that differences between potential outcome before treatment and potential outcome under the treatment are the same in the

[7] We have estimated robust standard errors using 1,000 stratified by sex bootstrap samples to avoid problems of heteroscedasticity.

[8] Besides the problem of the bias introduced by the failure of the assumption of parallel trends, the difference-in-differences estimator has a second problem related to the standard error of the estimates (Moulton 1990; Bertrand et al. 2004; Conley and Taber 2005; Donald and Lang 2007). We have used stratified by sex sampling bootstrap in order to estimate robust standard errors in the logistic regressions coefficients.

Table 3 Logistic regression estimates of parents on leave in the reference week (sample of men and women)

	β	Std[a]	Sig.	Exp(β)
Dependent variable leave (1 on leave; 0 not on leave)				
Treatment	0.0119	0.04313	0.791	1.012
Male	−3.6347**	0.10742	0.001	0.026
Treatment × male	0.9080**	0.12570	0.001	2.479
Constant	−0.6209**	0.03329	0.001	0.537
Number of observations	24,903			

−2 Log likelihood	16,469.371		
% Leaves = 1	15.7	χ^2 global adjustment	5,220.44
R^2 Cox and Snell	0.189	P value	0.000
R^2 Nagelkerke	0.325	Hosmer and Lemeshow	0.00
R^2 McFadden	0.241	P value	1.000

		Predicted			
		Leave = 0	Leave = 1		% Correctly classified
Observed	Leave = 0	14,397	6,584		68.62
	Leave = 1	360	3,562		90.82
Threshold 0.157				Total	72.12

	Estimated probabilities P(Leave = 1)	
Derivatives	Male (%)	Female (%)
Before treatment	1.4	35.0
After treatment	3.4	35.2
Absolute change ΔP	2.0	0.3
Relative change $\Delta P/P$	145.7	0.8

Sample 2005, 2006, 2008 and 2009 (2007 excluded), male and female wage earners with ages ranging from 16 to 64 and with children under 1 year

** Significant at 99 %

[a] Robust standard errors estimates using 1,000 stratified by sex bootstrap samples

Source EAPS

control group and in the treatment group (Athey and Imbens 2006). That is, this assumption assumes that changes in the outcome variable over time would have been exactly the same in both treatment and control groups in the absence of the intervention. The parallel-trend assumption is critical, and it implies that the DD estimator is appropriate when interventions are as good as random conditional on time and group fixed effects (Bertrand et al. 2004). The more similar, in terms of background characteristics, the treatment and control groups are, the more convincing is the DD approach (in the case of randomized experiment, treatment, and controls are identical for a large samples).

In our case, the DDs estimate is an unbiased estimate of the effect of the introduction of paternity leave if, in the absence of the policy reform on leave, the average changes in the difference in the take-up before and after 2007 would have been the same for fathers and mothers. The failure of the parallel trend assumption will occur when there are systematic differences between treatment and control groups, that is, between the behavior of fathers and mothers when taking leave. And precisely, as maintained by Han et al. (2009), the employment and leave-taking behavior of mothers and fathers is likely to differ markedly and may be differentially affected by leave policies.

The way to avoid failure of parallel trends is to have the most similar treatment and control groups as possible. In our case, where groups are not homogeneous, the best strategy to prevent the existence of parallel trends is to consider explicitly in the regression (1) all the covariates and factors (x_i) that could be explaining systematic differences between mothers' and fathers' take-ups in the absence of the treatment.[9] It is also convenient to perform simple differences between males and females along covariates to see whether they differ systematically. That is, we also introduce interactions between covariates and the male independent variable "male". At the same time, the estimate of these interactions will help us to explore in detail the determinants of fathers' use of childbirth leave (the second aim of this article). Thus, the DDs logistic regression conditional on covariates and factors that we use is:

$$\begin{aligned} P(\text{leave}=1) = \Phi[&\beta_0 + \beta_1\text{treatment} + \beta_2\text{male} + \beta_3\text{treatment}\times\text{male} \\ &+ \delta_1 x_1 + \gamma_1 x_1 \times \text{male} + \delta_2 x_2 + \gamma_2 x_2 \times \text{male} + \cdots], \end{aligned} \tag{2}$$

or alternatively

$$\begin{aligned} \text{Log}\left[\frac{P(\text{leave}=1)}{1-P(\text{leave}=1)}\right] = \beta_0 + \beta_1\text{treatment} + \beta_2\text{male} + \beta_3\text{treatment}\times\text{male} \\ + \delta_1 x_1 + \gamma_1 x_1 \times \text{male} + \delta_2 x_2 + \gamma_2 x_2 \times \text{male} + \cdots + u. \end{aligned}$$

The independent variables x_i refer to a series of personal and socio-labor characteristics which may have an influence on the person making the decision to take childbirth leave (we have justified these variables before). These independent variables (most of them dichotomous variables coded "1" = yes and "0" = no) are the following: "Public Sector": working in the public sector, which is assumed to have a positive effect on the take-up rate by men (and by women). "Temporary contract": having a temporary contract, that is assumed to have a negative impact on take-up rates by men (and women). "High education": having a university degree or

[9] On the other hand, DD estimates are more reliable when you compare outcomes just before and just after the policy change because the parallel trends assumption is more likely to hold over a short-time window. With a long-time window, many other things are likely to happen and confound the policy change effect. In our case, we use in our estimation 2 years before the reform and 2 years after. Moreover, in this period there have been no other policy measures that would affect periods of paternity leave directly. In this period there have been measures to support families on the birth of children, but those fell mainly on the mother, and obtaining them did not depend on whether or not the father was employed or if he took paternity leave or not.

similar, assumed to be a positive influence on men's take-up rate (and women's). "Manager": managers in private companies and in the public sector (groups "10", "11", "12", "13", and "14" of the Spanish adaptation of the International Standard Classification of Occupations, ISCO), which is assumed to have a negative effect on take-up by men (and women). "Encouraging regions": residence in the Basque Country, Navarre, La Rioja, Castilla-La Mancha, Castilla León, and Murcia, which is assumed to positively affect men's take-up (and women's). "Immigrant": having a foreign nationality of a developing country, which is assumed to have a negative influence on men's (and women) take-up. "Employed spouse": the spouse has a job, which is assumed to have a positive effect on men's take-up rate (and women's). Finally, we consider the possibility that the willingness of men to take leave initially increases with age, and then potentially declines, which means that we assume a negative influence of the covariate "age" and also a negative effect of the covariate "age squared".

Table 4 summarizes these independent variables, showing the total number of men and women and the percentage of men and women on leave included in each category. In the case of the variable "treatment", the percentage of men on leave in the reference week in the period after the introduction of the 13-day paternity leave (the treatment) was 3.44 %, whereas this figure was 1.40 % for the period prior 2007 (when the treatment did not yet exist). This result is repeated for all men included in each of the 10 categories considered. Moreover, although always showing take-up rates lower for men than for women, the percentages of both males and females who were on leave in the reference week are higher for those that work in the public sector, have higher education, have working spouses, and for those living in the "encouraging regions". On the contrary, the take-up rates are lower for those working on temporary contracts, those who are managers and those who are immigrants.

Among the covariates to be introduced in Eq. (2), we have also incorporated the "unemployment rate" in the region of residence of each individual. With this we want to take into account the possible effect that the economic crisis may have had on willingness to take childbirth leave. In a context of high insecurity of employment it may be that more males (and females) were reluctant to demand their rights and ask for childbirth leave (whether in the form of paternity leave, transferable part of maternity leave, unpaid parental leave, etc.), because of the fear of a possible adverse reaction from the employers.

Finally, we have also included as an independent variable in the regression model (2), a "trend" variable as control for different trends in men and women. With these variables we try to capture any changes not reflected in the above covariates that may have affected their willingness to use childbirth leaves.

We show in Table 5 the results of the estimate of the logistic regression (2). The global fit of the regression is satisfactory. The likelihood ratio test does reject the global significance of all independent variable coefficients. The Hosmer–Lemeshow test, that assesses whether or not the observed event rates match expected event rates in subgroups by using the χ^2 distribution, indicates that the model as a whole has sufficient explanatory power (*P* value greater than 0.05). Overall, the model explains 72.12 % of cases (predicted vs. observed).

Table 4 Percentage of males and females who were on leave in the reference week, disaggregated in accordance with eight determinants of parents' use of childbirth leave

	N		Concentration (%)			% Men on leave over total male wage earners in each group			% Women on leave over total female wage earners in each group		
						Total (%)	Treatment		Total (%)	Treatment (%)	
	Men	Women	Men	Women	% Women		Before (%)	After (%)		Before (%)	After (%)
Public sector											
Yes	2,241	2,431	15.2	24.0	52.0	4.06	2.37	5.80	36.98	35.20	38.77
No	12,516	7,715	84.8	76.0	38.1	2.15	1.22	3.03	34.52	34.87	34.25
Temporary contract											
Yes	3,713	2,381	25.2	23.5	39.1	1.91	1.66	2.21	25.83	24.62	27.02
No	1,1044	7,765	74.8	76.5	41.3	2.62	1.30	3.80	37.95	38.59	37.47
High education											
Yes	5,462	5,784	37.0	57.0	51.4	2.69	1.60	3.66	36.88	36.02	37.55
No	9,295	4,362	63.0	43.0	31.9	2.29	1.29	3.30	32.76	33.62	32.04
Age											
Age 16–21	104	177	0.70	1.74	63.0	2.88	1.67	4.55	16.95	13.58	19.79
Age 22–49	14,544	9,953	98.56	98.10	43.3	2.37	1.30	3.38	35.40	35.32	35.46
Age 50–64	109	–	0.74	–	–	11.93	11.59	12.50	–	–	–
Immigrant											
Yes	1,604	789	10.9	7.8	33.0	1.31	0.85	1.66	25.48	25.18	25.64
No	13,153	9,357	89.1	92.2	41.6	2.58	1.46	3.68	35.92	35.60	36.19
Employed spouse											
Yes	7,888	8,747	53.5	86.2	52.6	2.99	1.61	4.20	35.99	35.72	36.22
No	6,869	1,399	46.5	13.8	16.9	1.81	1.18	2.47	29.59	29.12	29.87
Manager											

Table 4 continued

	N		Concentration (%)		% Women	% Men on leave over total male wage earners in each group			% Women on leave over total female wage earners in each group		
						Total (%)	Treatment		Total (%)	Treatment (%)	
	Men	Women	Men	Women			Before (%)	After (%)		Before (%)	After (%)
Yes	509	230	3.4	2.3	31.1	1.96	2.22	1.76	33.04	31.65	33.77
No	14,248	9,916	96.6	97.7	41.0	2.46	1.37	3.50	35.16	35.02	35.27
Encouraging regions											
Yes	4,109	2,732	27.8	26.9	39.9	2.94	1.78	3.95	37.77	37.88	37.69
No	10,648	7,414	72.2	73.1	41.0	2.24	1.26	3.22	34.12	33.94	34.28
Total	14,757	10,146	–	–	40.7	2.44	1.40	3.44	35.11	34.96	35.23

Sample 2005, 2006, 2008, and 2009 (2007 excluded), male and female wage earners with ages ranging from 16 to 64 and with children under 1 year

Source EAPS

The variable "treatment" (the existence of paternity leave from the second quarter of 2007) has no statistically significant effect on the probability of the individual interviewed being on leave in the reference week. This result is normal, since there are far more women than men on leave, and they are not affected by the introduction of the paternity leave. Only when we consider the interaction between the variable "treatment" and the variable "male", a positive, statistically significant result is obtained. That is, following the introduction of paternity leave (the treatment), there is a greater probability that the men (treatment group) are on leave in the reference week. Thus, we have confirmation of the Hypothesis 1 that in Spain the introduction of paternity leave has increased the extent to which childbirth leave is being used by men.

On the other hand, and unlike the results of regression (1), after taking into account all other covariates, the explanatory variable "male" by itself is no longer significant. This confirms that by introducing these covariates and their interactions with "male", we must be properly reflecting systematic differences between mothers and fathers. In fact, our results show separate response for fathers and mothers to the same explanatory variables, reflected in the significance of the interaction between each covariate and "male".

With respect to the explanatory variables, in the first place, working in the *public sector* has a positive and statistically significant effect on the probability of being on leave, in both men and women. However, the positive effect on males for working in the public sector is higher than in women. In the public sector, with greater stability, better working conditions and a more family-friendly environment, men are more encouraged to request childbirth leave (Hypothesis 2). Moreover, the fact that, in Spain, in some public administrations men are entitled to more than a fortnight of paternity leave may influence the result. For example, male employees of the Catalonia Generalitat (regional administration for Catalonia) or the Madrid City Council are entitled to 4 weeks' leave for paternity leave.

Working on a *temporary contract* has a negative and statistically significant effect for men (Hypothesis 2) and for women. In Spain, the percentage of wage earners with temporary contracts is very high (in the sample used in this work, temporary contracts accounted for 25.2 % of the total male and 23.5 % of total female). Among such workers there are high levels of insecurity of employment, and this may lead to some of them not exercising their right to childbirth leave because of fear of their employers' reaction. This negative effect of having a temporary contract, however, is significantly lower in men than in women.

The results obtained for the variables "age" and "age squared" are quite interesting. Both variables are statistically significant but only for males. Thus, while age does not appear to exert any effect on the probability of being on leave in the case of women, there is a clear and significant U-shaped effect of age on probability of being on leave for men. In fact, the sign obtained for the variable "age squared" is the opposite of that originally proposed for male (Hypothesis 3). Indeed, we previously hypothesized a non-linear relationship, with the willingness of fathers to take leave initially increasing with age, and then potentially declining. However, in our estimate "age $\times$ male" has a negative effect, and "age squared $\times$ male" a positive one on the probability that the male is on leave in the reference week. This means that both the youngest and oldest fathers, with children under 1 year of age,

Table 5 Logistic regression estimates of parents on leave in the reference week (sample of men and women) controlling for differences in covariables

	β	Std[a]	Sig.	Exp (β)
Dependent variable leave (1 on leave; 0 not on leave)				
Treatment	0.010	0.120	0.946	1.010
Male	1.923	1.293	0.129	6.839
Treatment × male	1.066**	0.341	0.002	2.903
Public sector	0.196**	0.054	0.001	1.217
Public sector × male	0.398**	0.141	0.007	1.489
Temporary contract	−0.577**	0.057	0.001	0.562
Temporary contract × male	0.420**	0.150	0.005	1.523
High education	0.071	0.049	0.142	1.074
High education × male	−0.115	0.130	0.374	0.892
Age	0.022	0.041	0.588	1.022
Age squared	0.000	0.001	0.441	1.000
Age × male	−0.342**	0.072	0.001	0.711
Age squared × male	0.005**	0.001	0.001	1.005
Immigrant	−0.281**	0.087	0.001	0.755
Immigrant × male	−0.295	0.263	0.238	0.744
Employed spouse	0.155*	0.067	0.022	1.167
Employed spouse × male	0.291*	0.129	0.026	1.338
Manager	−0.187	0.147	0.203	0.829
Manager × male	−0.086	0.398	0.811	0.917
Encouraging regions	0.140**	0.048	0.005	1.151
Encouraging regions × male	0.119	0.126	0.337	1.127
Unemployment rate	−0.260	0.470	0.577	0.771
Unemployment rate × male	−0.727	1.331	0.577	0.484
Trend	0.001	0.010	0.911	1.001
Trend × male	−0.010	0.028	0.702	0.990
Constant	−0.914	0.660	0.159	0.401
Number of observations				24,903
% Leaves = 1				15.7
R^2 Cox and Snell				0.198
R^2 Nagelkerke				0.340
R^2 McFadden				0.253
−2 Log likelihood				16,203.402
χ^2 Global adjustment				5,486.41
P value				0.000
Hosmer and Lemeshow				6.49
P value				0.592

Table 5 continued

		Predicted			% Correctly classified
		Leave = 0	Leave = 1		
Observed	Leave = 0	14,388	6,593		68.58
	Leave = 1	358	3,564		90.87
Threshold: 0.157				Total	72.12

Derivatives:[b] Factor *j*	Male (%)	Female (%)
Treatment (after vs. before)[b, c, d]	142.4	0.7
Public sector(vs. private)[b, c, d]	63.2	13.7
Temporary contract (vs. fixed)[c, d]	−12.5	−33.7
High education (vs. low education)	−3.6	4.9
Immigrant (vs. national)[c]	−39.9	−17.7
Employed spouse (vs. not employed) [c, d]	46.2	11.0
Manager (vs. not manager)	−21.2	−12.1
Encouraging regions (vs. rest of regions)[c]	24.6	9.7

Sample 2005, 2006, 2008, and 2009 (2007 excluded), male and female wage earners with ages ranging from 16 to 64 and with children under 1 year

** Significant at 99 %; * significant at 95 %

[a] Robust standard errors estimates using 1,000 stratified by sex bootstrap samples

[b] *Derivatives* Estimated mean growth rate on the probability of leave when the factor j is changed: $P(\text{leave} = 1|\bar{x}'\boldsymbol{\beta},\ x_j = 1)/P(\text{leave} = 1|\bar{x}'\boldsymbol{\beta},\ x_j = 0) - 1$

[c] The factor is significant

[d] Significant differences between male and female in this factor

Source EAPS

have a higher probability of being on leave than the rest. This was already covered in Table 4. There it can be seen that in the group of very young working fathers, aged between 16 and 21 years, 2.88 % of them were on leave; while for the group aged 22–49 years (the middle section including most men in the sample) this figure fell to 2.37 %; and yet, for the older group of working fathers, aged 50–64 years, the percentage that was on leave amounted to 11.93 %, a high figure compared with the others. This result could be interpreted in different ways. Firstly, among men aged 50–64 years, traditional gender attitudes may have more influence (on average), which would affect their take-ups negatively. However, the members of this group have (on average) a professional and economic situation which is more stable and consolidated, and this facilitates access to leave. Additionally, since for this group to have a child is not very common, it may be that for some of these "old men" the arrival of a new baby is "something exceptional", and may give rise to highly motivated new fathers living the experience and showing quite advanced parenting behavior. This may be translated into a greater desire to take childbirth leave. And secondly, males aged 16–21, being very young, may be least likely to be in an employment situation providing secure access to leave arrangements, and this could negatively affect their take-ups. However, this effect can be offsetting in three ways: the youngest may have gender attitudes which are rather advanced and predispose them to take leave. Since their careers are at a very early stage, they sacrifice few opportunities when they take leave (low opportunity cost of taking leave). And finally, in Spain to be entitled to paternity leave is not very difficult (at least 180 days of social security contributions).

Note that the negative sign of the variable "age × male" indicates that for the central interval of ages the hypothesis that age negatively affects the male take-up seems to be confirmed (Hypothesis 3).

The variable "high education" has no statistically significant effect on the probability of being on leave in the reference week, either for males or for females. The same thing happens with the variable "manager", that has a negative effect on the probability of a male and female interviewee being on leave, but it is not statistically significant. Our variable "manager" corresponds to the major group "1" of the Spanish adaptation of the International Standard Classification of Occupations (ISCO-08). This is a fairly broad definition of a manager, with a high degree of heterogeneity (it includes senior government officials, corporate managers, department managers, managers of small business, etc.). Had it been possible to disaggregate this variable further, it would have been interesting to consider just the top managers, and check whether in this case a negative and significant relationship was obtained.

Being an "immigrant" (not from advanced economies) seems to have a significant negative influence on the probability of the male and female interviewee being on leave during the reference week, not having a significant differential effect between men and women. This last means that probably the negative effect of higher levels of job insecurity among immigrants (males and females) has more explanatory power than the negative effect of a more traditional gender attitudes among immigrants (that would affect only to the take-up rate of males).

The variable "employed spouse" has a statistically significant and positive effect on the probability of the male and female interviewee being on leave during the reference week. Nevertheless, this effect is much more important in the case of men than women. This result confirms Hypothesis 6 that the willingness to take leave increases among those males whose partners are working, compared to those whose partners are housewives.

In the "encouraging regions" the leave system is slightly more extensive than in the rest of regions, and, in some cases (Basque Country), gender equality policies encourage relatively the involvement of fathers in caring for small children. Thus, it may be expected that in these regions men (and women) will have greater take-up rates. The finding obtained here confirms this (Hypothesis 8), since the variable "encouraging regions" has a positive and statistically significant effect on male and females take-up rates, although no significantly different effect among men and women is detected (Lapuerta 2013 also obtained evidence of this positive effect among women, but not among men).

Neither the general "trend" variable nor its interaction with "male" are significant. This result is showing that when the number of periods is small, as in our case (2 years before and 2 years after the policy change), it is difficult to capture the trend in the changes in attitudes and behavior among parents in relation to childbirth leaves. This result supports our view that the parallel trend assumption remains valid in our model after controlling for the other independent variables.

Finally, the effect of the regional unemployment rate on the probability of being on leave does not appear to be significant. It seems that, at least with data from 2008 and 2009, high unemployment rates have not changed take-up rates of fathers and mothers, although the employment crisis has continued and intensified since 2009, and maybe this result could have been changed, and be significant if we were to perform regression analysis using a sample of later years.

Further Checks on the Assumption of Parallel Trends

To conclude our empirical analysis we perform a series of final checks, common in this type of studies (Meyer 1995), to corroborate that with the introduction of the covariates and their interactions in the logistic regression (2), we have eliminated the problem of failure in the assumption of parallel trends, and that our DDs estimator is unbiased. These final checks consist of the estimation of the effects of some "placebo" natural experiments with our sample.[10]

In Table 6, we show the results of the estimation of eight different models for four placebo experiments. In the first one we have used the same dependent and independent variables that were used in model (2), but only for the subsample of years 2005 and 2006, and we have used 2006 as a fictitious treatment. In the second model, we have changed the subsample to the years 2008 and 2009 using that last 2009 as a fictitious treatment. In both models, and since in these years there has not been any measure that would affect our response variable $P(\text{leave} = 1)$, our DDs estimator of the policy effect (interaction treatment $\times$ male) should be non-significant.

The third and fourth "placebo" experiments uses the complete sample 2005, 2006, 2008, and 2009, and the years 2008 and 2009 as "treatment", as in the original model, but we have replaced our dependent variable $P(\text{leave} = 1)$ by other response variables that are not supposed to be affected by the reform in paternity leave. Firstly, we have used as dependent variable the total number of children, more specifically $P(\text{more than one child} = 1)$. Secondly, we have also used as dependent variable the number of hours in a usual working-week, and more specifically $P(\text{usual working-week over 40 h} = 1)$. None of these dependent variables have, in principle, been affected by the introduction of paternity leave; therefore, the DD estimator of the effect of treatment on these dependent variables should be zero.

As shown in Table 6, in none of the four placebo experiments significant DD estimates (interaction treatment·male) were obtained, even when we control for

[10] In this DDs regression framework, it is also useful to throw covariates interacted with the "Treatment" dummy to control for changes in the composition of controls and treatment groups before and after the introduction of paternity leaves and thus to avoid the problem of parallel trends (Meyer 1995). We have tried to include these interactions in the model (2) (in addition to the interaction between "male" and covariates), but only the *Public Sector by Treatment* interaction has been weakly significant (only at 93 % confidence), so finally we have not included these interactions in our regression.

Table 6 Logistic regression estimates for placebo natural experiments

Models	Sample years 2005 and 2006 with placebo treatment in 2006		Sample 2008 and 2009 with placebo treatment in 2009		Sample 2005, 2006, 2008, and 2009		Sample 2005, 2006, 2008, and 2009	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Dependent variables	On leave		On leave		More than one child		Usual workweek over 40 h	
Treatment	−0.025	−0.040	−0.006	−0.151	0.089	0.047	−0.073	0.167
	[0.060]	[0.126]	[0.055]	[0.114]	[0.040]	[0.120]	[0.074]	[0.209]
	(0.669)	(0.746)	(0.918)	(0.190)	(0.029)	(0.698)	(0.328)	(0.416)
Male	−3.646	5.811	−2.713	1.895	0.194	−2.932	1.400	−3.186
	[0.151]	[2.222]	[0.097]	[2.143]	[0.039]	[1.280]	[0.060]	[1.012]
	(0.001)	(0.006)	(0.001)	(0.372)	(0.001)	(0.022)	(0.001)	(0.003)
Treatment × male	0.021	−0.464	−0.028	0.495	−0.024	−0.030	−0.029	−0.323
	[0.212]	[0.445]	[0.139]	[0.304]	[0.052]	[0.158]	[0.081]	[0.240]
	(0.918)	(0.279)	(0.858)	(0.117)	(0.645)	(0.853)	(0.721)	(0.174)
Covariates[a]	–	146.531	–	168.84	–	3,014.544	–	937.749
		df 22		df 22		df 22		df 22
		(0.000)		(0.000)		(0.000)		(0.000)
Constant	−0.608	−2.063	−0.606	−0.349	−0.239	−9.519	−2.403	2.445
	[0.044]	[1.095]	[0.040]	[1.026]	[0.030]	[1.121]	[0.053]	[0.893]
	(0.001)	(0.049)	(0.001)	(0.733)	(0.001)	(0.001)	(0.001)	(0.008)
Number of observations	11,755		13,148		24,903		24,903	
% Dependent variable = 1	14.3		17.0		47.9		18.6	
R^2 Cox and Snell	0.208	0.217	0.170	0.181	0.002	0.116	0.054	0.089
R^2 Nagelkerke	0.370	0.388	0.284	0.302	0.003	0.155	0.088	0.145
−2 Log likelihood	6,931.6	6,785.0	9,537.6	9,368.7	34,422.2	31,407.7	22,506.1	21,568.3
χ^2 Global adjustment	2,734.3	2,880.8	2,453.1	2,621.9	56.0	3,070.5	1,392.8	2,330.6

Table 6 continued

Models	Sample years 2005 and 2006 with placebo treatment in 2006		Sample 2008 and 2009 with placebo treatment in 2009		Sample 2005, 2006, 2008, and 2009		Sample 2005, 2006, 2008, and 2009	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
P value	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000
% Total correctly classified	74.1	74.1	70.4	70.5	51.7	64.9	52.8	59.5

Each column show: coefficient beta estimated; [robust standard errors]; and (*P* value for bilateral significance). Robust standard errors estimates using 1,000 stratified by sex bootstrap samples. Sample: male and female wage earners with ages range from 16 to 64 and with children under 1 year

[a] Likelihood ratio test (degrees of freedom, *P* value χ^2 statistic) to test the global significant of the covariables introduced in the regression: public sector; public sector × male; temporary contract; temporary contract × male; high education; high education × male; age; age × male; age squared; age squared × male; immigrant; immigrant × male; employed spouse; employed spouse × male; manager; manager × male; encouraging regions; encouraging regions × male; unemployment rate; unemployment rate × male; trend; trend × male

Source EAPS

systematic differences between men and women and the covariates used in the model (2) were introduced in these placebo experiments. These results reinforce the validity of our assumption of parallel trends after controlling for observables, and therefore we can conclude that our estimates of model (2) are unbiased.

Conclusions

In March 2007, the first paternity leave, non-transferable and of 13 days' duration, was introduced into Spain. It was supposed to be an intermediate stage with a final aim of attaining a month-long paternity leave. In fact, this month-long permission was intended to be introduced in January 2011, but the Spanish Government has delayed the introduction of this leave by at least a year as a result of the economic crisis. The 13-day paternity leave is very limited when compared with those existing in other countries such as Sweden, Germany, or Iceland. Despite all this, it is interesting to evaluate to what extent the introduction of this new leave has encouraged Spanish men to take up childbirth leave to a greater extent once they have become fathers.

For this purpose, a sample from the EAPS, comprising the 20 quarters which elapsed between 2005 and 2009, was used, so it has been possible to ascertain what number of male and female wage earners with children under 1 year were on childbirth leave in the reference week, in each of the quarters analyzed. From a methodological viewpoint, the introduction of a fortnight's paternity leave in the first quarter of 2007 also provides a case of a natural experiment. Fathers of children born immediately after the reform are "treated" differently to fathers of children born immediately before the reform.

The women–men gap in the use of leave has lessened as a consequence of the introduction of paternity leave. The rate at which this leave is used by men (the percentage of males on leave in the reference week compared to total number of men and women on leave during that week), for the period prior to the measure, was 5.99 %, while the subsequent rate of use was 11.34 %. Furthermore, the percentage of males on leave in the reference week, in the period prior to the first quarter of 2007, was 1.40 %, while that figure stood at 3.28 % for the period after the first quarter of 2007.

However, to make a rigorously empirical test of the hypothesis that introducing paternity leave has encouraged greater take-up by men, we proceeded to carry out a DDs estimation, with the sample of fathers and mothers 2 years before the reform (2005 and 2006) and 2 years after (2008 and 2009). We have conditioned the logistic regression on covariates and factors explaining the systematic differences between male and female concerning the probability of being on leave to avoid the failure on parallel trend assumption.

The results show a positive and statistically significant coefficient for the interaction between the variable "treatment" and the variable "male". This is interpreted in the sense that after the introduction of paternity leave (the treatment), there was an increased probability that the males (treatment group) would be on leave during the reference week. These findings confirm the hypothesis that in Spain the introduction of paternity leave has increased the degree to which men use the Spanish childbirth leave system.

Besides, regarding the determinants of the fathers' use of childbirth leave, this article has found that the probability of being on childbirth leave is higher among those fathers working in the public sector (this positive effect also occurs among mothers, but with less intensity), those with employed spouse (this effect also occurs among mothers, but with less intensity), and those living in the "encouraging" regions (there is no significant differential effect among men and women). On the other hand, the probability of being on childbirth leave is lower among fathers on temporary contracts (this negative effect also occurs among mothers, but with more intensity) and among fathers who are economic immigrants (this effect is not significantly different in the case of mothers). The age affects negatively the take up of fathers (not so in the case of mothers) but, interestingly, the oldest fathers (with children under 1 year) of the sample have a higher probability of being on leave. For the older group of fathers, aged 50–64 years, the percentage that was on leave in the reference week amounted to 11.93 % (the average percentage for the total sample was 2.47 %).

With respect to policy implications, Ridgeway and Correll (2004) argue that widely shared, hegemonic cultural beliefs about gender and their impact on social relational contexts are among the core components that maintain and change the gender system. And they conclude that the gender system will only be undermined through the long-term, persistent accumulation of everyday challenges to the system resulting from socioeconomic change and individual resistance. A fundamental line, which challenges the system, is the one occurring through the ever-increasing number of men doing domestic work and caring for children. And one of the main lines of action in public equality policies is that of accelerating this trend.

One basic way of speeding up the trend toward males caring for small children is to make access to childbirth leaves easier for them. In the case of Spain, the introduction of the 13-day paternity leave is a public policy which has increased the extent to which men use childbirth leave. However, the woman–man gap in the use of childbirth leave is still high. One of the causes of this difference is the short duration of Spanish paternity leave. For this reason, extending it to a month's paternity leave (as foreseen by the government), is a measure which, undoubtedly, would contribute to reducing that difference. However, this should be no more than a transitional stage toward the goal of achieving complete gender equality in Spain with regard to access to leave, in line with the equal and partly non-transferable leaves existing in Sweden, Iceland, or Germany (Castro and Pazos 2008).

But even assuming the achievement of full legal gender equality in the access to childbirth leaves, policy should encourage fathers to take greater advantage of leave arrangements, via a series of complementary measures, such as the following: (a) helping parents reconcile work and family life, either indirectly, by encouraging firms to apply flexible workplace practices or family-friendly arrangements; or directly, for example by providing good quality formal childcare support, (b) promoting equality policies that incorporate co-responsibility targets (egalitarian distribution of family responsibilities between the couple) by, for example, increasing the duration of paid leave entitlements that are non-transferable between parents; increasing information to both parents about fathers' rights to parental leave (OECD 2007, pp. 21–22); and increasing information about fathers' rights to use reconciliation practices (Escot et al. 2012).

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