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DISCUSSION PAPER SERIES

IZA DP No. 13756

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SEPTEMBER 2020

ISSN: 2365-9793

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ABSTRACT

Work Disability after Motherhood and How Paternity Leave Can Help*

We study how childbirth increases the likelihood of young, working mothers to claim disability insurance and how paternity leave could ease this effect. Our event study analysis uses Belgian data to show that the incidence rate of disability across gender only diverges after first-time childbirth. This "other child penalty" can be reduced with the provision of paternity leave. Our regression discontinuity difference-in-differences design shows that mothers with partners eligible for a two-week-long paternity leave spent on average 21% fewer days on disability over twelve years. Moreover, we show links between this incidence of paternity leave and consequent birth-spacing decisions.

JEL Classification: J16, J13, I13, H55

Keywords: disability insurance, gender, child penalty, paternity leave, maternal health, birth spacing, natural experiment, regression discontinuity, event study

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* This research project benefited from the financial support of the Thiepolam fund. We thank seminar participants at the Belgian Day for Labour Economists (Maastricht, 2019), IEB workshop Social Expenditure (Barcelona, 2019), ULB Centre Emile Bernheim (2019), DULBEA workshop on the Economics of Disability (2019), Luxembourg Institute of Socio-Economic Research (2019), SOFI Workshop Gender Economics (Stockholm, 2019), Louisiana State University (2019), IWEPS (2019), Université catholique de Louvain (2019), University of Nevada at Reno (2019), IAE Workshop Health & Labour Market (Nantes, 2019), XXVII Meeting Public Economics (Barcelona, 2020), National Bank of Belgium (2020), University of Groningen (2020), Erasmus University Health Economics seminar (Rotterdam, 2020), EALE-SOLE-AASLE World Conference (2020) for comments and suggestions. We also thank Thomas Barnay, Patrick Button, David De la Croix, Delia Furtado, Paula Gobbi, Camille Landais, Erica Lindahl, Olivier Marie, Noemi Peter, François Rycx, Ity Shurtz, Todd Sorensen, Ariane Szafarz, Judit Vall Castello and Philippe Van Kerm for valuable comments.

1 Introduction

It is well documented that motherhood is particularly detrimental to a woman's professional career. Recent studies highlight a child penalty for a mother's earnings of between 12% and 55%, depending on the country considered. Similar dramatic differences are not observed for men. In fact, the fraction of gender wage inequality caused by "child penalties" represent today, in most countries, the main factor behind gender labor market inequalities (Angrist & Evans, 1998; Bertrand, 2011; Kleven, Landais, & Søgaard, 2019; Kleven, Landais, Posch, Steinhauer, & Zweimüller, 2019; Lundborg, Plug, & Rasmussen, 2017; Neumeier, Sørensen, & Webber, 2018). What is not well documented, however, is another child penalty that may prove to be equally important. And so, in this paper we study how parenthood could also trigger gender differences in disability and how paternity leave could offset this potential imbalance. We find that, up to eight years after childbirth, mothers are 40% more likely to experience work disability than fathers. Our research demonstrates that while having children increases a mother's probability to enter disability, it does not seem to affect fathers. We show next that paternity leave mitigates the time mothers spent on disability by 21% over a period of 12 years after childbirth. This effect is more pronounced for mothers who have their first child below the age of 30 and for individuals suffering from musculoskeletal disorders. Lastly, we provide evidence that links the introduction of paternity leave to an increase in birth spacing, which in turn negatively impacts the likelihood of women to enter disability after childbirth.

Our study contributes to a better understanding of gender inequalities in the context of Disability Insurance (DI). This is of particular importance given that the number of persons deemed unable to work for health reasons and receiving DI benefits has increased substantially in OECD countries, particularly among women, creating an important challenge for social security funding (OECD, 2010). With that aspect in mind, we add to the literature on DI by providing new insights about the causes contriuting to work disability for fathers and mothers at young ages (Autor & Duggan, 2006; Dahl & Gielen, 2020; Liebman, 2015).

To estimate the "other" child penalty that links motherhood with an increased proclivity towards disability, we use an event study approach similar to the one used by Kleven, Landais & Søgaard (2019). This approach is based on individual-level variations in the timing of first births and sharp changes that occur around childbirth. Our analysis using Belgian administrative data reveals a disability-specific child penalty that does not disappear over the long run and, even up to eight years after their first child's birth, mothers are 1.2 percentage points more likely than fathers to enter DI. This represents a child penalty for women of around 40%. We also demonstrate that the impact of children increases with the size of the family, with a gender gap that reaches 2.3 percentage points for parents with three children. While postpartum health effects have been highlighted before (Cheng, Fowles, & Walker, 2006; Saurel-Cubizolles, Romito, Lelong, & Ancel, 2000), we are among the first to capture the long-term health consequences that take place so many years after childbirth. We believe that this result provides significant new insights into women's career trajectories and the specific role that breaks due to sickness and disability play on their labor market attachment.

Our findings are consistent with two recent studies on Sweden and Norway (Andresen & Nix, 2019; Angelov, Johansson, & Lindahl, 2020). In the former, Angelov *et al.* (2020) rely on within-couple variations to show that mothers more than double their sick leave compared to fathers after the birth of their first child. Our study differs because our event study approach allows us to also look at mothers and fathers separately. In their study on Norway, Andresen and Nix (2019) interestingly look at child penalties after birth for both heterosexual and lesbian couples. While their findings for mothers in heterosexual relationship are similar to ours, the ones for lesbian couples do not show postnatal differences in sickness absences between mothers who bore the child and the one who did not. This result gives support to our argument that long-term effects may be driven by family arrangements rather than the biological cost of giving birth. On the downside, all their findings rely on a measure of sickness that also includes absence for dependents, including young children. Consequently, they cannot accurately disentangle direct health effects on the mother from days-off used to take care of young children. In contrast, our study relies on disability spells that have been validated by a doctor and concerns only the health of the mother. This allows us to capture in our findings any direct health effect without any other interference.

We next argue that the "other" child penalty we observe for mothers, or their increased probability to enter disability, might be linked to family arrangements which require employed women to work a "second shift" at home (Hochschild & Machung, 1990) and take on a larger share of domestic work, including child care. This is well-documented in Guryan, Hurst, & Kearney (2008) who have put together time use surveys from 14 countries and have showed that, while the gender gap in time spent with children varies across countries, it is always detrimental to women. The ratio of childcare hours between mothers and fathers ranges from 2 to 1 in countries like Canada, the Netherlands, Norway, the United

States or Belgium,[1] while it exceeds 3 to 1 in Austria or France. Moreover, this across-country imbalance is often combined with another reality whereby working mothers spend less time devoted to childfree leisure and personal care (Craig, 2007; Parker & Wang, 2013; Pepin, Sayer, & Casper, 2018).[2] Taken together, this combination of more domestic work and less leisure for working mothers might ultimately affect women's health and career, and explain a higher likelihood for them to suffer from work disability.

Building on the initial findings of this study, we next look at whether the provision of a paternity leave could be an effective policy to moderate a mother's entry into disability. Our reasoning draws on Becker's (1985) theoretical framework that associates a woman's career with her household responsibilities and relies on numerous studies that have empirically shown how paternity leave policies effectively increase a father's involvement in child care, even when they are short (Farré & González, 2019; Hook, 2010; Kotsadam & Finseraas, 2011; Patnaik, 2019; Tamm, 2019).[3] Taken together, it suggests that if the inequality of family arrangements and a mother's tendency to fall into disability are linked, then paternity leave provisions could impact mother's health, thus softening this "other" child penalty. In the short run, that would translate into more help from the father right after birth. More critically in the long run, it could affect permanently the division of tasks in households, with long-lasting consequences for women. We exploit a discontinuity in Belgian legislation which opened a two-week paternity leave only to fathers of children born after July 1st 2002 to analyze its effect on the probability for women to enter into disability after childbirth. To do so, we use a regression discontinuity difference-in-differences (RD-DiD) framework similar to Avdic & Karimi (2018). This research design relies on the exogeneity of the time of birth to compare households that had a child before and after the July 2002 reform. It also considers using non-reform years to wash out any seasonality in disability through its difference-in-differences dimension (Avdic & Karimi, 2018; Cygan-Rehm, Kuehnle, & Riphahn, 2018; Danzer & Lavy, 2017; Farré & González, 2019; Lalive, Schlosser, Steinhauer, & Zweimüller, 2014). We implement this research design to a sample of around 100,000 children born between 2002 and 2004 for which we have administrative data on their parental labor market history, including disability spells and benefits

[1] For Belgium, Table A1 in the Appendix shows that working mothers spent on average more than double the time on childcare than their partners in 2013.

[2] See Table A1 in the Appendix for a similar pattern in the Belgian case.

[3] A complete balance between men and women might, however, not be achieved. For example, Ekberg, Eriksson & Friebel (2013) show that fathers affected by a paternity leave reform in Sweden did not take more days-off to take care of sick children.

received.

Overall, we find that the introduction of a two-week paternity leave decreases the number of days on DI for women by 21% up to 12 years after childbirth. The decrease in the number of days in disability is the largest for individuals on DI for more than 12 months. In that case, the decrease is indeed equivalent to a 33% reduction. This result highlights the benefit of paternity leave for women's career since individuals on DI for more than 12 months have on average a much lower probability to re-enter the labor market. Results on the impact of the paternity leave reform on DI benefits confirm the different findings on the number of days. By contrast, we find no evidence of any change in days or benefits for fathers that are eligible for a paternity leave. If so, the positive effect on mother's health does not seem to be at the expense of father's, suggesting an overall positive health's impact of paternity leave at the household level. All these results are robust to alternative specifications including varying the bandwidth selection or the trend definition and to a series of placebo tests.

We next present evidence that our main results are entirely driven by an effect on mothers who had their first child during the reform year. For this specific group, the number of days on disability up to 12 years after childbirth were 40 percent lower. Contrarily, women experiencing the birth of additional children after their first child in 2002 did not experience any change after the introduction of the two-week paternity leave. This striking contrast could provide suggestive evidence that, in eligible families, first-time parents are less imbedded in fixed roles and are thus more inclined to change their behavior when the father takes a paternity leave (Patnaik, 2019; Sundström & Duvander, 2002). Focusing finally on the causes of disability, we show that 50% of the long-term reduction in the number of days on DI happens for mothers with musculoskeletal disorders. These disorders represent up to 40% of all disability cases for mother in the long-run (Saurel-Cubizolles et al., 2000).

Our results relate to a growing literature on the effects of paternity leave policies. An important part of this literature focuses on its causal impact on fathers' and mothers' labor supply and wages. While a series of these studies find a positive effect on women's earnings and their probability to participate in the labor force (Andersen, 2018; Druedahl, Ejrnæs, & Jørgensen, 2019; Dunatchik & Özcan, 2019; Farré & González, 2019; Rege & Solli, 2013), others do not reach similar conclusive results (Cools, Fiva, & Kirkebøen, 2015; Ekberg et al., 2013).[4] Few

[4]Our paper is also connected to recent studies that have reported effects of paternity leave on outcomes like divorces which is not directly related to the labor market but might still

studies, however, focus like us on the effect of paternity leave on mother's health and disability spells (Persson & Rossin-Slater, 2019; Ugreninov, 2013). And when they do, they only consider short term health effects, while we observe up to 12 years of disability status after childbirth. Ugreninov (2013) focuses on Norway and, contrary to us, does not find any significant effect of paternity leave on mother's health. A possible reason for this difference in the result is that she does not take seasonality in her outcome variables into account like us and the most recent papers on paternity leave.[5] The other study that tackles mother's health focuses more on the impact of a greater degree of flexibility in taking a paternity leave rather than a net effect like we do (Persson & Rossin-Slater, 2019).[6] Using an RD-DiD design like the one we use, Persson & Rossin-Slater (2019) find that increasing a father's leave flexibility reduces a mother's risk of physical postpartum health complications and improves their mental health within the first 6 months after childbirth.

Our paper is also related to the few studies analyzing fertility decisions in conjunction with parental leave policy for fathers (Cools et al., 2015; Duvander, Lappegard, & Johansson, 2020; Farré & González, 2019). In this study, we consider changes in fertility patterns from the perspective of the role these changes can play in explaining the effect of paternity leave on disability. In this light, we show that our main results are entirely driven by first-time mothers who are younger than 30 years-old and still fertile. For this group, the number of days on disability up to 12 years after childbirth was about 47% lower when their partners were eligible for paternity leave. Second, we show that the paternity leave reform in Belgium has increased birth spacing between the first two children for those same first-time young mothers. We argue that these results could provide suggestive evidence on the role of spacing in explaining the reduction of days on DI for mothers after the reform of the paternity leave system in 2002. We make the case that both results exhibit similar time dynamics and are driven by the

affect a woman's career. Avdic & Karimi (2018) show, for example, that couples who were affected by the introduction of a paternity leave quota in 1995 in Sweden, the so-called "daddy-month", increased their probability of separation compared to unaffected couples. This result differs, however, from the finding of Olafsson and Steingrimsdottir (2020), who show that the introduction of one month of parental leave earmarked to fathers increased marital stability in Iceland.

[5]Another explanation could be related to the fact that Ugreninov (2013) does not use a "classic" regression discontinuity design and ends up for that reason with a very small sample of parents who had a child within one month of the reform date.

[6]The paper evaluates a particular Swedish policy called "double days' that allows fathers to take up to 30 days of paid leave on an intermittent basis alongside the mother during the first year of the child, without affecting total leave duration

same sub-population, that is younger mothers who could decide to delay the birth of their second child. We conclude, therefore, that the timing of births in families with multiple-children is key to reducing the problem of work disability of mothers who have children before they turn thirty.

The remainder of the paper proceeds as follows. Section 2 provides information on the institutional settings of the disability insurance system and the parental leave policy in Belgium. Section 3 focuses on the event analysis. Section 4 aims at presenting the regression discontinuity difference-in-differences analysis. Section 5 concludes.

2 Institutional Context

2.1 Parental Leave

The Belgian parental leave system has gradually developed since the seventies.[7] In 1971, a new law provided for the introduction of a 15-week[8] paid maternity leave around childbirth. This program remains in effect at the time of writing and combines both pre- and post-birth leave with the obligation to be off work at least one week before birth and 10 weeks overall. Prospective mothers qualify for this paid leave if they have worked at least 120 days[9] in the last 6 months.[10] The replacement rate is 82% of their gross salary during the first 30 days (uncapped) and 75% thereafter (capped at a ceiling of 2810 euros per month as of January 1, 2020).

The parental leave system was expanded in July 2002 with the introduction of a comprehensive, job-protected, paid paternity leave for fathers with a salaried contract.[11] Before the introduction of this law, fathers of newly born children were only entitled to 3 days of paid job absence.[12] The new paternity leave program

[7] Table A2 in the appendix reports its main features.

[8] 19 weeks for multiple births.

[9] 400 hours if they work part-time.

[10] The unemployed mothers are also eligible for the same program if they have 120 active days of job search in the last 6 months before birth. Civil servants are entitled to the same program in the same conditions but with a different benefits system. Finally, the self-employed mothers are eligible for a different program paid at a flat rate and offering twelve weeks of maternity leave.

[11] This program includes all private sector workers and contractual employees in the public sector. Similar programs exist for civil servants but are directly managed by the different public administration. A completely separate paid paternity leave program for the self-employed was introduced in 2019.

[12] Since 1978, fathers were allowed to take 3 days off work after the birth of a child, called “congé de circonstance”, which is equivalent to specific leaves for attending weddings or funerals

introduced an additional period of 7 working days, which together with the 3 days of job absence, brought the leave period to 2 weeks. Initially, fathers had to take their paternity leave during the first month after childbirth, but the time frame was extended to 4 months in 2009, hence allowing fathers to take their paternity leave after the compulsory maternity leave period of mothers. As for the replacement rate, the first 3 days are fully compensated by the employer, while the remaining 7 working days are compensated like the mothers at 82% of the gross salary. As shown in Figure A1, a substantial number of fathers opted into this policy after it was introduced during the second half of 2002, and kept increasing in the following years.[13]

It might be useful to put the Belgian system in perspective with other countries, notably the Scandinavian countries, which were early adopters of government paid leave policies accessible to the fathers. In Sweden, for instance, the parental leave system was introduced in 1974 and was gender neutral. Both the mother and the father were given an equal amount of paid leave for their children, but with the option of freely transferring paid leave days between each other. The system was reformed in 1995 to encourage fathers to take a bigger share of the parental leave. A so-called "daddy-month" was introduced, reserving 1 month of paid leave to each parent, implying that 1 month of paid leave would be lost if either parent chose not to take any leave.

In Belgium, parental leave has never been transferable between parents. We believe that this feature makes it a particularly interesting case for research, since fathers can take paternity leave without an automatic reduction for mothers. In other words, we can measure the net effect of providing paternity leave. Many studies in Scandinavian countries actually measure the combined effect of the paternity leave provision and the reduction of maternal leave (e.g. Avdic & Karimi, 2018; Ekberg et al., 2013). In some cases, the reforms are even combined with an increase of the total leave period for both parents (e.g. the second "daddy month" reform in Sweden), which makes it even harder to disentangle the estimated effects. In the context of work disability, the reduction of maternity leave could

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offered by the "loi du 3 juillet 1978 relative aux contrats de travail". This category of leave is not fully job-protected and thus not like the paternity leave introduced in 2002.

[13]Finally, on top of the specific maternity and paternity leave programs described above, parents are also individually entitled to 4 months of parental leave that they can take at any time before their children turn 12 years old. The leave can be taken simultaneously by both parents and can also be taken over 8 months for a career interruption of 50% or 20 months for a career interruption of 20%. Workers who decide to use this form of leave receive a fixed amount instead of a percentage of their salary, which could make it less appealing in many cases. In 2017, 70% of the beneficiaries of this program were women (IEFH, 2018).

have detrimental effects on maternal health, which might not be balanced by the provision of paternity leave. This could be another reason why Ugreninov (2013) does not find any effect of paternity leave provision on mothers' sick leave absence in Norway.

2.2 The Belgian Disability Insurance System

In this paper, we focus on the health of workers by observing how much time they spend on disability benefits. In Belgium, employed workers with a minimum number of working days have access to disability benefits through the National Institute for Health and Disability Insurance (NIHDI).[14] It covers them against health shocks that affect their ability to work for at least one month.[15] The application terms and conditions vary, however, between disability spells that are either less than a year and those that are longer.[16] In the remainder of the paper, we will therefore distinguish between these two types by referring respectively to the "short-term disability" spells and the "long-term disability" ones.[17]

In order to qualify for short-term disability coverage, individuals must be recognized as "unable to work" by a doctor designated by their health insurance fund.[18]A worker would be considered eligible when his/her ability to work is reduced by at least 66% with respect to the last occupation.[19] To qualify, the applicant should also have stopped all productive activity as a consequence of a deterioration of his/her health that is not directly related to his/her professional activity.[20] If these two conditions are still applicable after a year, a disabled worker may qualify for long-term disability status. There is, however, no automatic tran-

[14]Full-time workers and unemployed workers must have fulfilled a minimum of 180 working days (or active days of job search for the unemployed) during the last twelve months to be eligible. For part-time workers, the condition is to have worked at least 800 hours over the last 12 months.

[15]The same insurance also applies to the unemployed. The self-employed have access to a distinct disability insurance program that we do not study here.

[16]Spells shorter than a month are fully paid by the employers and are not covered by this insurance program.

[17]The disability literature sometimes refers to temporary and permanent disability programs.

[18]In Belgium, although the health care system is publicly supported at the national level, the reimbursement of medical expenses and short-term disability benefits are paid through the public health insurance funds called "mutualities", which are funded by the NIHDI and act as intermediaries with the disabled. In short, to benefit from the Belgian medical coverage, individuals must register at a health insurance fund.

[19]Note that an important change occurs after 6 months of disability: the reduction in the ability to work is then evaluated with respect to any occupation that the worker could perform given his/her age, education and experience (instead of his/her previous occupation).

[20]This condition exists to establish a distinction between the disability insurance program and other programs such as the occupational injuries fund and the occupational diseases fund

sition from the short-term status to the long-term one. In order to be accepted into the long-term disability program, the applicants' doctor (who oversaw the applicant during the short-term period) has to submit the application to the NIHDI, which can either directly approve the doctor's decision or run its own internal evaluation.

The replacement rate also varies with the duration of the disability spell. During the first year it amounts to 60% of the last wage payment received before becoming disabled.[21] After one year, when one enters the long-term disability program, the replacement rate depends on the last wage payment received,[22] as well as the position of the disabled person in the household. To be precise, this share is 65% for heads of households, 60% for single households and 40% for cohabitants, with defined floor and ceiling amounts.[23]

Figure A2 (Panel A) in the appendix shows the evolution of the disability rate for the working-age population in Belgium since 1980.[24] Like in many OECD countries, the number of persons receiving DI benefits has increased substantially and particularly among female beneficiaries. It is often argued that this high increase for women reflects, in part, their growing labor force participation, which contributed an expansion of the pool of insured workers, as more and more women had sufficient work history to qualify for DI. But according to Autor and Duggan (2006), this would explain only about one-sixth of the increase in the rate of female DI beneficiaries. Consistent with this, Figure A3, which considers eligible workers only, shows that the incidence rate for women is growing faster than for their male counterparts. This is true for Belgium (Panel A), the origin of data used in this study, but also for the United-States (Panel B) and most OECD countries (OECD, 2010).

Another important trend in DI results from reforms[25] that expanded the eligibility criteria and induced major changes in the composition of the beneficiary population, with a notable shift towards younger workers. Autor and Duggan (2006) explain that these new legislations place more weight on "applicants' reported pain and discomfort", making it easier to qualify for certain impairments

[21]For unemployed, it is equal to the unemployment benefits.

[22]For unemployed workers, it is the last wage payment before unemployment.

[23]In 2020, the maximum short-term disability benefits were 2,248 euros per month, while the maximum long-term disability benefits were 2,435 euros per month.

[24]We plot DI beneficiaries from the long-term program, which is more directly comparable to the U.S. Social Security Disability Insurance (SSDI) program, often studied in the DI literature (e.g Autor & Duggan, 2006; Liebman, 2015).

[25]In the United-States, the major reform was implemented in 1984 by the Disability Benefits Reform Act (P.L. 98-460).

that used to be "hard to verify", such as back pain or depression (Liebman, 2015). The side effect of these reforms has been an increased incidence rate of disability at younger ages (Congressional Research Service, 2018). Indeed, mental and musculoskeletal disorders tend to have an early onset and low age-specific mortality (Autor & Duggan, 2006). As a result, those beneficiaries are likely to enter early on the DI program and experience a relatively long duration. In 2017, 65.7% of the Belgians on long-term disability benefits were suffering from mental and musculoskeletal disorders.[26]

Hence, while work disability used to concern mostly older men prior to the 1990s, it is now increasingly affecting women, and particularly at younger ages. Our study adds to the existing literature on DI by exploring these gender inequalities among young adults. We show that this gender inequality as related to disability insurance can by partially explained by parenthood and by how couples react to the arrival of children in the household.

3 Event Study Analysis of Work Disability after Motherhood

Our first research question evaluates to what extent children can affect the probability of their parents to fall into disability. As explained by Kleven, Landais & Søgaard (2019) the ideal experiment to do so would be to randomize fertility. In the absence of such an experiment,[27] they propose instead an event study approach based on individual-level variations in the timing of the birth of the first child to capture its direct effects on different labor market outcomes. The rationale being that, although fertility choices and the timing of birth are not exogenous, the outcomes of interest should evolve smoothly over time. Thus, any sharp changes around childbirth are likely to be orthogonal to unobserved determinants and seize any causal effects (Kleven, Landais, & Søgaard, 2019). In our case, it might be argued for example that women who invested in their education are more likely to have children later in life and are less likely overall to enter disability. However, the effect of education on those outcomes should not generate any sharp changes and therefore should be disregarded as an explanation linking childbirth to parental disability.

[26] Administrative data from NIHDI:
https://www.riziv.fgov.be/fr/statistiques/indemnites/Pages/default.aspx

[27] Lundborg, Plug and Würtz Rasmussen (2017) have come up with another very convincing strategy that uses *in vitro* fertilization treatments.

The event study approach has the additional advantage of tracing out the full dynamic trajectory of the effects over time, therefore capturing the impact of the first child, as well as of any subsequent children. Previous studies using instruments for the number of children, such as twin births (Bronars & Grogger, 1994; Rosenzweig & Wolpin, 1980) or the gender breakdown of siblings (Angrist & Evans, 1998), could only succeed in estimating local effects of second or higher order children. Our approach will instead capture the overall impact of having children on the probability to enter DI for mothers relative to fathers.

Event studies have been used in different contexts, those regarding the impacts of inheritances (Druedahl & Martinello, 2016), hospital admissions (Dobkin, Finkelstein, Kluender, & Notowidigdo, 2018) or family health shocks (Fadlon & Nielsen, 2019). In our specific setting, we foresee one limitation, the fact that this framework will not allow us to measure the impact of choices made before parents had children. For instance, if women invest less in education and career in anticipation of motherhood, then the estimated child penalties represent the lower bounds on the total lifetime impacts of children (Kleven, Landais, & Søgaard, 2019). In other words, our study will be able to identify the post-child effects of children conditional on choices made before parenthood.

3.1 Data & Empirical Strategy

We use a rich set of administrative data from the Belgian Crossroads Bank for Social Security (CBSS) to conduct our different empirical analyses. This database puts together several administrative registers linked at the individual level and contains quarterly information on social security status over time, household composition and labor market history. Importantly for our research design, the data allows us to match children with their parents through the National Registry and to observe the exact month of childbirth. Regarding data on disability, we can observe the disability status during any given quarter, as well as the number of days of each disability spell and the amount of benefits received. As part of this study, we obtained a large sample of 60% of all births during the years 2002 to 2013, with stratification at the provincial level to ensure representativity. From this sample, we were able to identify the parents and build a dataset that tracks their disability status quarterly over the period from 2002 to 2016.[28]

For the event study analysis, we narrow our sample to all individuals who had their first child between 2002 and 2013, without imposing any restrictions on

[28]This corresponds to a sample of 861,344 births and 1,271,079 parents.

the relationship status of the parents. This leaves us with an estimation sample of 691,922 parents, including 359,657 first-time mothers and 332,265 first-time fathers. We follow those parents over a period of up to 12 years, including up to 4 years prior to the birth of their firstborn and up to 8 years after it.[29] In total, we observe each parent during 48 quarters.

We now turn to the econometric setting of the event study analysis. For each individual in the data, we first denote by t = 0 the quarter-year in which the father/mother has his/her first child and index all quarters relative to that time period. We then analyze changes in the disability status as a function of event time both in the short- and long-term, estimating the following equation separately for men and women:

$$y^g_{iqt} = \sum_{j \neq -4} \beta^g_j \cdot I[j = t] + \sum_k \gamma^g_k \cdot I[k = age_{iq}] + \sum_y \delta^g_y \cdot I[y = q] + \epsilon^g_{iqt} \tag{1}$$

where y^g_{iqt}, our main outcome of interest, is a dummy variable to indicate the receipt of disability benefits during a given quarter q for individual i of gender g and at event time t. On the right side, equation (1) includes a full set of event time dummies (first term on the right-hand side), age dummies (second term) and time period dummies (third term). We omit the event time dummy at t = -4, implying that the event time coefficients β^g_j measure the impact of children relative to four quarters before the first child's birth. We voluntarily chose a date not too close to childbirth, as we suspect that short-term disability would raise for women during their pregnancy. Following Kleven, Landais & Søgaard (2019), we include a full set of age dummies to control non-parametrically for underlying life-cycle trends. Additionally, the age dummies improve the comparison between men and women, as women are on average a few years younger than men when they have their first child. In addition, we include a full set of quarter-year dummies to control non-parametrically for time trends and seasonal effects. Finally, we also control for linear pre-trends to consider potential pre-childbirth differences between men and women that age and quarter-year dummies would not capture and that could bookend the breaks around parenthood. To do so, we follow Kleven, Landais & Søgaard (2019) and estimate a linear trend separately for men and women using

[29] Our sample includes parents who had a child between January 2002 and December 2013. We follow those parents until 2016. Our panel is therefore unbalanced because the follow-up period differs according to the birth date of the reference child. For parents who had a child in 2002, we do not have data on the four years before. For parents who had a child in 2013, we have pre-birth outcomes but a reduced follow-up period of 3 years. We ran the estimations on a perfectly balanced panel and found similar results (available on request).

only pre-event data (i.e. from quarter -16 to quarter -4 before birth), and then use the result in the main event study specification described in equation (1) to residualize the outcome variable with the estimated pre-trend.

3.2 Main Results

Figure 1 (Panel A) plots the gender-specific impacts of children on disability status across event time. The outcome includes both short and long-term disability.[30] As explained above, it corresponds to changes in disability rates at event time t relative to the 4^{th} quarter before the first child's birth (t = -4), having controlled non-parametrically for age and time trends. The figure also includes 95% confidence bands around the event coefficients. Several lessons could be drawn from Figure 1 (Panel A) regarding parenthood, disability and how their interaction could impact men and women differently.

First, we know from our data that the disability rate is equal to 2.8% for both men and women at t = -4, and so there does not seem to be any gender difference in the disability rate before the birth of a first child. From there, however, the situation changes dramatically for women but not that much for men. Indeed, women experience a sharp increase in their probability to enter disability starting 3 quarters before their first child's birth. The timing corresponds to the beginning of the pregnancy and reflects in most of the cases pregnancy-related health issues. This sharp increase peaks in the quarter right before childbirth with an increase of about 6 percentage points in comparison to the 4^{th} quarter before giving birth.[31] From there, the next three quarters show a gradual return to the pre-pregnancy level and to a situation in which both men and women seem to experience the same probability to be on disability benefits. This downward trend around childbirth is to a large part mechanical since all women, sick and eligible for disability benefits or not, slide to compulsory maternity leave for at least 9 weeks after delivery.[32]

Moving now to the results between one year and up to eight years after birth, we

[30]The data for short-term disability insurance in 2002 is available only for the four (out of six) biggest health insurance funds. As a robustness check, we also estimate the event study analysis excluding 2002. Results do not change (available upon request).

[31]The effect might be even larger given that women who are sick during the last six weeks of their pregnancy are already covered by their maternity leave and cannot be registered as disabled for that reason.

[32]As we can see in Figure 1, there are still women on disability in the quarter of childbirth and the next one. This situation is because women who are sick during the last six weeks before childbirth are only entitled to 9 weeks of postpartum maternity leave instead of 15. Consequently, a woman who gave birth at the beginning of a given quarter might still enter into disability during the same quarter if she was only entitled to 9 weeks of maternity leave after birth.

can learn from Figure 1 (Panel A) that women start to experience another increase of disability during the second and third year following their first child's birth.[33] After the third year, this change stabilizes and eight years after delivery reaches an increase of about 2.1 percentage points compared to the pre-birth disability rate. Bearing in mind that our event study design captures the total effect of all children, this last result suggests the existence of an overall long-term "disability penalty" from having children that only impacts women. In contrast, we also show in Figure 1 (Panel A) that men seem to be largely unaffected by children. We only detect a small increase in their probability to enter disability two quarters after their first child's birth that stabilizes itself at 0.9 percentage points, eight years after it. Most importantly, we observe that the probability of men and women to enter disability insurance never converges back and that eight years after their first child's birth, a 1.2 percentage points gap remains. Since the average disability rate at t-4 was 2.8% for both women and men, this corresponds to a child penalty for women that amounts to 43%. This finding suggests the existence of another extremely important "child penalty", in addition to the "child penalty" generally associated with income and job loss in the literature.

For comparison purposes, we also offer original estimates for these "classic" child penalties in Belgium. Figure A4 in the Appendix reports the impact of children on labor earnings, participation rates and hours worked for men and women in Belgium. Panel A in the figure shows a 43% drop in female earnings up to 8 years after childbirth. We do not observe a similar negative effect for fathers who do not experience any change in their earnings before and after the birth of their first child. This 43% reduction in income for mothers place the Belgian child penalty at the level observed recently in the UK (Kleven, Landais, Posch, et al., 2019) and between the penalties observed in the Scandinavian countries (around 25%) and the one experienced in German-speaking countries (around 55%) (Kleven, Landais, Posch, et al., 2019). The results presented in Panel B and C of Figure A4 complete this analysis by showing again an additional child penalty for mothers in participation rate and hours worked. As with earnings, fathers do not experience any similar negative drops in these measurements before and after the birth of a child. All in all, the birth of a first child seems to play a crucial role in the career of women through the negative effect it has on their earnings and labor supply but also on their long-term health. Nothing similar is to be found for fathers.

[33] Interestingly, this second increase coincides with the average time of the arrival of subsequent children in an household. We would come back to that dynamic in the next sections.

3.3 Heterogeneous Effects by Household Size

In this subsection, we want to observe how the effects measured in the event study analysis could vary with the total number of children in a given household. Even though our event study is based on parents who had their first child between 2002-2013, the results presented in Figure 1 (Panel A) are based on the full sample, irrespective of the total number of children they end up having. As already explained, this means that the dynamics we observe include the effects of children born after the first one. In other words, the estimated long-run impacts should be interpreted as capturing the total effect of all children. To explore the implications of multiple children, we replicate the event study analysis on 3 subsamples that we split based on the total number of children which parents produce - 1, 2, and 3 children, respectively as of 2016.[34]

Figure 1 also presents the results of this analysis by the overall number of children in a household. The figure shows that the sharp increase around the birth of the first child is roughly similar in magnitude for the three subsamples. We also notice that the coefficient for mothers reverts to a level close to zero in the third quarter after childbirth for all types of families. It is only from the fourth quarter after childbirth that trends start differing across households. In families with a single child (Panel B), the trends between parents are only slightly different. The gender gap eight years after the birth of their only child reaches only 0.8 percentage point. In families with two children (Panel C), we observe an increasing gap between mothers and fathers in the second and third year following the birth of the first child. This very likely captures the effect of the second child. The gap between mothers and fathers up to eight years after the birth of their first child reaches 1.4 percentage points. It is expected that the two-child families (Panel C) look very much like the estimates for the whole sample in Panel A, since those families make up 50% of our sample. Finally, in families with three children (Panel D), the gap between parents reaches 2.3 percentage points after eight years. Placed end to end, these findings strengthen our conclusions that the probability for women to enter disability depends strongly on having children and increases when they have more than one. These findings also reiterate that dynamic has little to no impact on fathers and their probability to be on disability benefits.

One might ask, however, whether the increased probability for women to enter DI reflects merely the multiple pregnancies and deliveries or corresponds to the larger cost of having and providing for multiple children. To answer this concern,

[34] It gives us three samples of 31%, 50%, and 15%of all the women included in our main analysis. The remaining 4% have more than 3 children.

we replicate our event study analysis around the second child's birth, conditioning our sample on having two children in total, as of 2016. From Figure A5 in Appendix, we observe a spike in the probability of women to enter DI around the second childbirth that is similar in magnitude to Figure 1 for the first child. We also see a small bump during the four years that precede the second child's birth. This is of course related to the first child's birth. It is a smooth bump rather than a sharp spike because the birth of the first child did not take place during the same quarter for all women. More interestingly, we note, is the increase in DI probability that follows the second child's birth. Since we conditioned our sample on households with two children, this subsequent increase cannot be attributed to other childbirths. We believe that it instead reflects the long-run effects for women of having multiple children. Thus, we conclude that beyond the short-term effects related to giving birth, there are indeed long-term health effects of having children for women, which are reflected in their increased probability to enter DI even eight years after their second child's birth.

4 Paternity Leave and Maternal Disability

In the previous section, we provide empirical evidence that children have a large impact on the probability of mothers to enter disability. We now turn to study whether paternity leave could be an effective policy to moderate the entry of women into disability after motherhood. Interested by both the short- and long-term consequences of the policy on women, we focus our analysis on the cumulative effects over a period of up to 12 years after childbirth. Within this framework, we also try to capture any tradeoff the policy could create for fathers and analyze how family planning decisions could play a role in this context.

4.1 Empirical Strategy

We use a regression discontinuity design to analyze the impact of paternity leave on maternal disability, exploiting a cutoff in the Belgian legislation, which opened paternity leave only to fathers who had a child after the 1st of July 2002. It relies on the fact that fathers, whose children were born right before that date, did not have access to this newly introduced two-week leave. We implement the method on a sample of parents who had a child in a 6-month window around the reform.[35].

[35]In Section 4.5, we test the sensitivity of our results to different bandwidth selection, i.e. incrementally changing the window's size from 6 months to 1 month around July 2002, the

Our running variable is the month of birth. As explained by Imbens (2008), the key assumption of this design is that individuals are unable to manipulate the assignment variable. In our case, this seems like a reasonable assumption since birth dates are arguably difficult to manipulate. If this assumption holds, having a child right before or right after July 1st is as good as random.

All our specifications estimate intent-to-treat (ITT) effects since we observe eligibility (month of birth) but not the actual take-up of paternity leave. Indeed, individual-level data on paternity leave is not available for the second half of 2002, the year that the policy was introduced. We, however, have data on the subsequent years for children born in 2003 and 2004. In our sub-sample of parents who were both working at the time of birth, the take-up was respectively 55.6% in 2003 and 57.5% in 2004. Thus, our estimates suggest effects on the treated that are up to twice as large as our ITT estimates. Finally, to assure that we properly capture a causal effect of the paternity reform at the cutoff, we also need to assume that there are no other important changes of relevance (such as other policy interventions) for parents of children born right after the 1st of July. We are not aware of any such potentially confounding factors.

Taken together, these different elements motivate the estimation of the following regression-discontinuity design model:[36]

$$y_i^T = \alpha + 1[t_i \geq c]\beta + 1[t_i \geq c] \cdot f_r(t-c, \gamma_r) + 1[t_i < c] \cdot f_l(c-t, \gamma_l) + \zeta X_i + \epsilon_i \quad (2)$$

where y_i^T is the outcome of interest, T quarters after birth, for each parent of child i born in month t. c is the reform cutoff month, $1[\cdot]$ is the indicator function, f_l and f_r are unknown functions with parameter vectors γ_l and γ_r, capturing trends in the outcome of interest. We can interpret β as the estimated discontinuity for a given outcome when having children born just before and just after the 1st of July 2002. And if we assume that parents do not have exact control of when their children are born in a period around the 1st of July cutoff, we can interpret the estimated discontinuity as the causal effect of the paternity leave reform. Finally, we included a vector of control variables X_i, for age of parent, number of kids and region of living at the time of the birth of the reference child. We know that those variables might affect the probability of entering disability and should therefore help us get more precise estimates. We test formally at the end of this Section that those predetermined outcomes are perfectly balanced between the treatment

month of the reform

[36] See Lee and Lemieux (2010) for an in-depth description of the regression-discontinuity design econometric framework.

and control groups.

The only remaining issue concerns the potential seasonality of our main outcome variable. Figure A6 in Appendix shows that the number of disability days does indeed vary according to the moment that the child is born. We observe that women who had a child during the second part of the years 2003 and 2004 (non-reform years) always have on average a higher number of disability days. We therefore need to account for this seasonality when we measure the discontinuity in 2002. To do so, we combine the regression discontinuity design of equation (2) with a difference-in-differences model in a way similar to other research on the topic of parental leave (e.g. Avdic & Karimi, 2018; Cygan-Rehm et al., 2018; Danzer & Lavy, 2017; Dustmann & Schönberg, 2012; Farré & González, 2019; Lalive et al., 2014). We then apply this combined approach to the sample of children born during the reform year (2002) as well as during two non-reform years (2003 and 2004). This approach is valid under an additional common trends assumption that our outcomes' trends are comparable between reform and non-reform years. We cannot think of reasons why the seasonality pattern would change because of the introduction of the paternity leave.

This setting has the additional advantage of accounting for the fact that we capture the disability status in our data at the quarter level while our running variable is identified monthly. This mismatch could be problematic as it creates mechanically differences in the follow-up period between couples whose children were born at the beginning or at the end of a quarter. For instance, if we observe the outcomes of parents one quarter after birth: those who had a child in June exhibit follow-up periods ranging from 3 to 4 months, while parents of children born in July exhibit follow-up periods ranging from 5 to 6 months. This might be important since the discontinuity will be measured between June and July, which are respectively the end and the beginning of a quarter. Using a regression discontinuity difference-in-differences design will also help solve that problem by washing out any such mechanical correlation between the month of birth and the probability to be on disability benefits (Avdic & Karimi, 2018).

Specifically, we extend equation (2), using years 2003 and 2004 on top of 2002, add an indicator R={0, 1}, equal to one for the reform year 2002 and zero otherwise, and interact this new variable with each variable included in the model:

$$\begin{aligned} y_i^T = \alpha + \sum_{s=0}^{1} 1[R_i = s] \cdot \{1[t_i \geq c]\beta_s + 1[t_i \geq c] \cdot f_r(t - c, \gamma_{rs}) \\ +1[t_i < c] \cdot f_l(c - t, \gamma_{ls})\} + \zeta X_i + \lambda_n + \epsilon_i \end{aligned} \quad (3)$$

Equation (3) is essentially a fully interacted version of (2) with separate effects for reform and non-reform years, with the exception of fixed effects for each non-reform year, represented by λ_n. Our coefficient of interest is still β_1, which is now the interaction between "having a child after July 1st" and the 2002 indicator (R). By doing so, this new specification controls for systematic differences in outcome across families having a child in different (even if close) months of the year.

In terms of data, we restrict the sample used in the event study analysis to households with children born between January 2002 and December 2004 in which both parents are known at the time of birth but do not necessarily form a couple (in the sense of marriage or cohabitation) and might not even live together. Since we are primarily interested in the effects of paternity leave on mothers, we exclude mono-parental families from our analysis. We also restrict the sample to those households in which both parents were working at the time of birth. Since paternity leave is only available for salaried men, we do not want to include in our sample households in which the father was not working at the time of birth. This leaves us with an estimation sample of 101,735 households.[37]

Before moving onto to the presentation of the results, we test the validity of our identification strategy. As explained above, our design relies on when the paternity leave policy was introduced (July 1st 2002) and on the timing of childbirth around that date. Taken together, these two elements imply that being part of the treatment group or the control one is as good as random. First, we show that there is no evidence that parents were able to self-select into the new paternity leave system. If that would have been the case, it would have invalidated our identifying strategy. Manipulating the date for natural births is virtually impossible, but we want to rule out that planned cesarean sections or induced labor were not rescheduled in order for fathers to become eligible for the new paternity leave. To do so, we use data from Statbel, the Belgian statistical office, on the number of daily births in 2002. We start by providing graphical evidence in Figure A7 that the frequency of daily births had not been affected by the reform. We observe that there is no evidence of bunching around the threshold.[38] As a second step, we test

[37]For the estimations focusing on the fathers in the appendix, the number of observations is different and stands at 99,502. It is mainly due to the fact that for 2 percent of our sample we do not have information on the fathers for one or more of the three control variables used in the estimations (i.e. number of children, age and region, at the moment of birth of the reference child).

[38]We see on Figure A7 that there are always fewer births during the weekend, likely due to fewer scheduled deliveries. The day of the introduction of the policy, July 1st 2002, was a Monday. Mechanically, we observe that there are more births on that day than on the two previous days.

for sorting formally by estimating regressions of the form of equation (2), using as an outcome the log number of daily births.[39] Column 1 in Table A3 reports the results using a 7-day window around the threshold. Each subsequent column in the table increases the window by a week up to a specification with a 42-day-long window around the threshold. The coefficients for the different specifications are all small and statistically indistinguishable from zero. These results indicate that there is no discontinuity in the number of births around the threshold and therefore suggest that families did not manipulate the date of childbirth to become eligible for the new paternity leave policy.

Table 1 further confirms the plausibility of our identifying assumptions by testing for a discontinuity around the reform cutoff for a large array of characteristics of the parents: their region of living, size of household, number of children, whether the reference child is their first child, age, type of employment (i.e. blue/white collar worker, civil servant or self-employed), as well as daily wage. All of these terms are measured in the quarter of birth of the reference child.[40] We test that the parents' characteristics are balanced around the threshold by applying equation (3) to all those observable variables. The right panel of Table 1 reports the results of these regressions and shows that all coefficients are statistically indistinguishable from zero. They confirm that there is no evidence of discontinuity in the characteristics of the parents who had children right before or after the introduction of the paternity leave policy in July 2002.

4.2 Main Results

Now that we have presented our RD-DiD framework and established the validity of the design, we can turn to the results of our analysis on the causal impact of paternity leave on the disability status of mothers. Table 2 displays the treatment effects for three disability-related outcome variables in the 12 years following childbirth. The first outcome, reported in Panel A, displays results for the total number of days on DI since childbirth. It captures in this way the effects of the reform both at the extensive and intensive margins. The second outcome for which the results are displayed in Panel B focuses on the total amount of disability benefits received over the 12 years following childbirth and the 2002 reform. This cumu-

[39] We control for a linear trend in all but the first regressions. We also include dummies for the day of the week (i.e. Monday, Tuesday, ...).

[40] The wages are measured the quarter before the quarter of birth of the reference child. We had to limit the sample to a 3-month window (instead of 6 months), because our data starts in 2002. Therefore, we cannot observe outcomes the quarter before birth for those who had a child between Jan. and March 2002.

lative effect also captures an impact at both the extensive and intensive margins, while accounting for differences in daily allowances. The last outcome, reported in the table as "Ever on DI" (Panel C), displays results for the probability to have entered disability insurance at least once over the 12-year period that follows childbirth. It captures any effect of the reform that would take place at the extensive margin only. For each outcome, we present the overall effect, as well as the effect breakdown between the short-term and the long-term disability programs. Finally, Table 2 also displays the average of the outcome variables in order to give a sense of the size of the impact of the paternity leave reform.

Focusing first on the whole sample ("All mothers" in Table 2), we find statistically significant evidence that mothers who had a child with a father eligible for paternity leave spent on average 22 fewer days on DI in the 12 years following childbirth. Given an average of 105 days on disability, this represents a reduction of 21%. This result suggests an effect on the treated that may be twice as high, since we estimate an ITT effect with a 50% take-up rate. Interestingly, we observe that this effect is most pronounced for the number of days on the long-term disability program. On average, mothers in the treatment group spend 16.1 less days in the long-term disability program and 6.3 less days in the short-term one. This corresponds respectively to a 33% and 11% decrease compared with the baseline average of each group. This particular result more explicitly demonstrates the important role that paternity leave plays on a mother's career since the long-term disability program involves individuals who have been away from the labor market for a long period of time and who do not, in most cases, possess a work contract (De Brouwer & Tojerow, 2018). This suggests that the introduction of paternity leave could be particularly effective at decreasing the number of days in disability that are more consequential for the attachment of women to the labor market.[41]

In line with the results for the number of days, Panel B in Table 2 shows that mothers in the treatment group display an average decrease in disability benefits of the same magnitude (18%, 712 euros) compared with a baseline mean of 4049 euros. As with the number of days, the effect of the paternity leave reform appears much more concentrated on the long-term program with a reduction of around 30% on the benefits received in the 12 years that follow childbirth. Taken together,

[41] Another reason why the time spent in the short-term disability program proves less damaging relates to the fact that employers, at least in Belgium, cannot terminate an open-ended contract during the first six months of a disability leave period. In practice, many workers on short-term disability will therefore go back to the same employer when their health allows. On the other hand, most workers on long-term disability have been laid off and need to find another job when their disability status ends.

our results show that after childbirth mothers whose partners were eligible for a two-week paternity leave saw their health to be significantly less affected over a period of 12 years. This finding shows up through both the number of days spent on disability and the benefits received. Interestingly, this positive effect on the health of mothers does not occur at the expense of the fathers. Table A4 in the Appendix displays the effect of the paternity leave on fathers for our disability-related outcomes. It shows that fathers seem unaffected by the introduction of the new policy, whether it be for the number of days on DI or for benefits received in the 12 years following childbirth. In both cases, we do not observe a statistically significant change one way or another and conclude thus that the reform has not been detrimental to the working health of fathers.

Finally, Table 2 also displays in Panel C results for the probability to have entered disability insurance at least once over the 12-year period that follows childbirth. Reported in the table as "Ever on DI", this outcome allows us to capture any potential effect of the reform that would have taken place at the extensive margin. Unlike the previous results, it indicates no statistically significant change for mothers during the 12 years following childbirth. This seems to indicate that most of the effect of paternity leave concerns rather serious health issues and takes place at the intensive margin rather than the extensive one.

Table 2 displays the aggregate outcome for the cumulative number of days in DI over the 12 year period following childbirth. Our RD-DiD setup also allows us to capture the dynamic effects of the paternity leave reform by estimating equation (3) for each quarter from the birth of the reference child (t=0) to 12 years after childbirth (t=48). Focusing on short- and long-term disability separately, Figure 2 plots the treatment effects of these regressions for the number of days on DI by quarter. In both cases, the dynamic pattern highlights a decreasing trend in the number of days over time, consistent with our previous results that mothers were affected by the paternity leave. As time passes, the beneficial effect of paternity leave is reflected in a commensurate increase in the number of days not on disability. Regarding the timing of the effect in the two programs, we observe that the effect becomes significant and negative around 3 years after childbirth in the short-term program (Panel A) and around 5 years in the long-term one (Panel B). The discrepancy in the timing logically reflects the need to spend a year in the short-term disability program before having access to the long-term one.

4.3 Effect of Paternity Leave by Birth Order

This section investigates how the treatment effects that we have identified could vary as a function of the birth order of children. Table 2 shows estimates separately for first-time mothers, which represent 48% of our sample, and "experienced" mothers who birthed a second or higher order child when the paternity leave reform was introduced in 2002. Regardless of the selected outcome, the effect seems entirely concentrated on those mothers who had their first child during the reform year. Panel A of Table 2 shows that first-time mothers, whose partners had access to paternity leave, spent on average 39 fewer days on disability overall. This result represents a decrease of 40%, that is almost twice as large as the effect observed for the whole sample. In parallel, we do not observe any statistically significant effect of paternity leave on the number of days in disability for mothers who had additional child during the reform year. Figure 3 illustrates this difference between the two groups in a dynamic way for the short- and long-term disability programs together. While the negative effect on first-time mothers starts to be significant after 2 years and slowly builds over time to reach 39 days after 12 years, the effect remains close to zero over the whole period for experienced mothers. Panel B of Table 2 further confirms this result as it relates to disability benefits, indicating a significant decrease in the total amounts for first-time mothers and no effect for the other group of mothers. Interestingly, Table 2 also reports a statistically significant effect at the extensive margin by showing that the probability to enter the long-term disability program is 2 percentage points lower than in the control group. This result is again concentrated on first-time mothers.

Put together with the other results from Table 2, this result reinforces our finding that paternity leave seems to have generated important changes in the long run in households lacking childcare experience and in which the respective roles related to child management have not yet coalesced. Those results are consistent with previous findings showing that the division of labor becomes more gender-based only after the birth of the first child (Kleven, Landais, & Søgaard, 2019; Kleven, Landais, Posch, et al., 2019) and that first-time fathers respond more strongly to policy incentives since their views or habits about child-rearing are not yet anchored (Patnaik, 2019; Sundström & Duvander, 2002). We believe that those results might also reflect subsequent fertility decisions. Indeed, during the follow-up period covered by our dataset, first-time mothers need to decide whether they want additional children, and if this is the case when they would prefer to have them. In the final section of this paper, we deepen the scope of such an argument and look in more detail at how changes in fertility patterns might play a role in explaining the effect of paternity leave on maternal disability.

4.4 Effect of Paternity Leave by Medical Condition

Here we explore heterogeneous effects depending on the medical condition for which DI beneficiaries have obtained their status. We do so with a special focus on mental and musculoskeletal disorders, which account for respectively 37% and 24% of the number of days on disability registered in our sample. Hence, we estimate equation (3) separately to look at the impact of the 2002 reform on the number of days on disability benefits for (1) mental disorders, (2) musculoskeletal disorders and (3) other health issues.[42] Since we only know the medical condition of the beneficiaries once they are categorized as having long-term disability status, we only perform this analysis on this particular program.[43] Table 3 presents results for the three types of health disorders and, as a comparison, for the whole sample now with a follow-up period of 11 years. In sum, these results demonstrate that almost 50% of the reduction in long-term disability days for mothers is related to musculoskeletal disorders. This corresponds to a decrease of 5.7 days in the long-term disability program for the mothers with a child born after July 2002. By contrast, the table indicates no statistically significant change in the number of days on disability for mental health disorders or any other type of health issues. Figure A9 in the Appendix shows dynamic estimates for these different groups and corroborates this result. It shows moreover that long-term disability days for musculoskeletal disorders (Panel B) start decreasing as early as two years after the reference child's birth and slowly accumulate over time to reach a total of about 6 days by the end of the period.

We conclude from this heterogeneity analysis that the long-term reduction in the number of disability days for mothers is largely driven by a decrease in disability related to musculoskeletal disorders. This result is not surprising given that the prevalence of backaches or back pain in general remains high among new mothers, even after the first postpartum year (Saurel-Cubizolles et al., 2000). Thus, the introduction of paternity leave in Belgium in 2002 seems to have unleashed a significant and positive impact on a widespread maternal health problem that

[42]Information on medical condition is based on the International Statistical Classification of Diseases. Using the first 2 digits, we identify 17 categories for ICD-9 until 2015 that we group in 3 categories (1) mental disorders, (2) musculoskeletal disorders and (3) others.

[43]We also restrict the analysis to 11 years after childbirth because there was a change in the ICD classification in 2016 without any possibility to convert the data of that year to the previous classification system.

historically has led mothers to seek disability benefits. To interpret this result in another way, our study could reveal moral hazard amongst mothers on disability insurance, since musculoskeletal disorders are among the "hard to verify" impairments (Angelov et al., 2020; Liebman, 2015). In that case, paternity leave polices, which encourage fathers to be more involved in childcare, would decrease the occurrence of mothers who use DI to spend time with their children. While we cannot completely exclude this, the fact that we do not observe any effect on other self-reported ailments, like mental disorders, makes this explanation highly unlikely. The moral hazard argument is not attuned to this self-diagnostic specificity.

4.5 Robustness Checks

In this section, we provide robustness checks for our RD-DiD design. We first test the sensitivity of our main results to the bandwidth selection. For our main specification, we use a bandwidth of 6 months, which is the largest window available given that the reform we study took place in July 1st 2002 and that the period covered by our dataset starts in January of that year. As a robustness test, we vary the bandwidth from 6 months to 1 month around the threshold to observe how the coefficient for the causal effect of the reform would change. We also run a donut-hole specification, excluding births that took place one month before and after the cutoff, to confirm that parents did not manipulate the birth date. Table A5 in the Appendix displays the effects on maternal disability-related outcomes for these different specifications. We conclude from the table that our findings are robust to the choice of bandwidth.

Then, we investigate the sensitivity of our main results to different trend definitions to obtain an unbiased estimate of the discontinuity at the cutoff. Since we *a priori* do not know the functional forms of f_l and f_r in equation (3), we test for linear trend (our main specification), as well as quadratic and cubic trends. From Table A6 in the Appendix, we see that the reduction in disability days for mothers is very similar whether we use a linear or quadratic trend, respectively -22 and -21. When using a cubic trend, the reduction is larger and amounts to -51 days. Regarding disability benefits, the amount varies from -466 (quadratic) to -1541 (cubic). Altogether, our findings appear to be robust to the model specification.

Finally, following Avdic & Karimi (2018), we use non-reform years in a "randomization inference design" and perform placebo analyses shifting artificially the reform cutoff by one month at a time. We estimate a placebo intervention 43 times between January 2003 and July 2006 using our RD-DiD design defined in equation (3). We estimate effects on our main outcome, that is the cumulative number of days on disability for mothers, but we restrict the period to 10 years after childbirth, which is the maximum follow-up period in our sample for women who had a child in 2006. Figure A10 in the Appendix shows the distribution of point estimates from this procedure (Panel A) and the cumulative distribution of t-values from the series of regressions (Panel B) compared to a standard normal distribution. The point estimates from the placebo interventions are almost always higher than our estimated effect of -13.7 days (indicated by the doted vertical line) and, as expected, centered around zero ($\overline{\beta}_{placebo}$ = -0.1). Furthermore, we perform normality tests on the empirical distribution of the placebo coefficients (Skewness and kurtosis test), as well as the cumulative empirical distribution of the t-values (Kolmogorov-Smirnov test). Both tests cannot be rejected for any conventional significance level. All in all, these placebo tests also reinforce the robustness of our main findings.

4.6 Fertility Decisions and the Effect of Paternity Leave on Disability

In this section, we explore how changes in fertility patterns could explain our central conclusion, which links paternity leave to a reduction in the time that mothers spend on disability benefits after childbirth. We focus particularly on the role of subsequent children in explaining the paternity leave effect for three reasons. First, we know from the event study analysis that the probability to enter disability on the long-run was higher for women with more children, suggesting a link between the number of children and the consequences of motherhood on health. Two, again because of the event study analysis, we associate a significant portion of the overall increase in the likelihood to become disabled to a particular spike that transpires in the second and third years following the birth of the first child. We attributed this second increase to the arrival of more children in the household suggesting again an association between subsequent births and the long-term health of mothers. Finally, our analysis in the previous section highlights that paternity leave impacts maternal disability most significantly when it is taken after the first child. Women in this context are particular because they are in the middle of their fertility window and have the option to decide if and when they want a second child. Together, these findings point to a potential explanation that links the impact of paternity leave on disability to decisions made in the context of family planning. In what follows, we attempt to provide evidence documenting the prevalence of this mechanism.

And so, we analyze how the two-week paternity leave introduced in 2002 may have impacted birth spacing and family planning fertility decisions overall. We estimate our RD-DiD specification using a series of new outcomes indicating if a mother had a second child within 1, 2, 3, 6, and 12 years after the birth of 2002.[44] We complete this series of one-off indicators with an overall indicator measuring the total number of children in the household twelve years after the birth of the reference child. The entire analysis is carried out on a sample of first-time mothers whose reference child was their first child in 2002. We limit our sample to this group of mothers following the results in Section 4.3 showing that the effect of paternity leave on maternal disability is only driven by households from this particular type. Since age is a central factor in explaining fertility decision, we also add that dimension in this analysis by estimating our specification by age.

Table 4 displays the estimated effects for all the fertility-related outcomes broken down by the age of the mother at birth. The table also reports averages of outcome variables for comparison purposes. We do not find any statistically significant evidence that paternity leave changes family planning decision for mothers above the age of 30. We do find, on the other hand, that the probability to have a second child for mothers, aged less than 30 and who had a first child with a father eligible for paternity leave, is about 5 percentage points lower in the second and third year after the first child's birth. This represents a decrease of 13.3% and 9.4% in the probability to have a second child in the second and third year, respectively. These results suggest that mother who had a first child with a father eligible for paternity leave took longer to have another child. In other words, birth spacing between the two first children increased for treated women.

This effect of paternity leave on birth spacing does not seem, however, to engender any overall effect on the likelihood of having subsequent children. Indeed, 12 years after the birth of the reference child, the size of families in our study is not significantly different between the families that were eligible for paternity leave and the ones that were not. Figure 4 (Panel A) illustrates that dynamic with more details by showing those effects on a quarterly basis. We can clearly discern the statistically significant negative effects on fertility around the second and third years and the gradual convergence to zero afterward. Figure 4 (Panel

[44]While we can observe the first born for each mother in our dataset based on information from the Belgian register of births, we do not observe exactly the subsequent births from the same mother but rather the year her household had one more child. It could thus also result from adoption or family recomposition.

B) also confirms that the subsequent fertility of mothers aged more than 30 years old is unaffected by the introduction of the paternity leave. We believe that since those mothers are closer to the end of their fertile cycle, they cannot easily adjust their birth spacing to the effects of paternity leave on the household. Overall, our results echo those of Farré & González (2019) for Spain. They found that a similar reform, that is the introduction of a two-week paternity leave, led to delays in subsequent fertility (Farré & González, 2019).[45]

We argue that the increase in birth spacing propelled by the introduction of the paternity leave could be the main mechanism explaining our results related to disability. This assumption rose from the close match in the timing of the effect of paternity leave on birth spacing and on the number of days on disability. If we go back to Figure 3 in Section 4.3, we can clearly see that the number of disability days for first-time mothers (Panel A) starts diverging from zero two years after the birth of the reference child, which is also exactly when birth spacing occurs. We thus think that this change in the number of days on disability could be driven by the delay in the birth of a subsequent child, as the dynamic clearly matches the one in Figure 4 (Panel A). This mechanism is all the more plausible given that our main results were driven by first-time mothers in the middle of their fertility window and who could thus consider increasing the time between their first two children.

Interestingly, these new results also put forward the importance of age in understanding the dynamics of paternity leave. For this reason, we went back to our main results concerning the effect on disability and separated first-time mothers below the age of 30 from first-time mothers above the age of 30 (respectively 61% and 39% of the "first-timer" sample). Reported in Table A7 in Appendix, all the results seem to indicate that the introduction of the paternity leave particularly affected the disability status of younger first-time mothers (below the age of 30 in 2002), while not affecting the older ones at all. At the intensive margin, it corresponds to a negative effect for these younger mothers that amounts to 44.5 fewer days on disability and 1687 euros less benefits, over a period of 12 years (both statistically significant at a 5% level). Those results suggest again that the same population is driving both the results on fertility and disability.

It could be argued, however, that this explanation of our results on the link between paternity leave and disability are driven by the fact that mothers who

[45] Farré & González (2019), however, found that older mothers had fewer children on average, while we find that their total fertility is unaffected. On the same subject but in the different context, two studies focusing on the Nordic countries did not find an effect on fertility following paternity leave reforms (Cools et al., 2015; Duvander et al., 2020).

delay their subsequent fertility are observed during a longer period of time with a single child since the twelve years follow-up period is indexed on the birth of the first child. In that case, if the event of having a second child increases the probability to enter disability and if this event is delayed, this would mechanically reduce the overall number of days on disability benefits and call into question our whole line of reasoning. To study this conjecture, we estimate the effect of the 2002 paternity leave reform on disability-related outcomes from the birth of the second child. In this new setting, the follow-up period is centered around the birth of the second child and all outcome variables capture disability spells that took place after that birth.[46] By doing so, we can wash out any mechanical effect due to a difference in the period covered after the second birth. In the same vein, we narrow also the period covered to measure the effect of the reform to 8 years after the birth of the second child to assure that both treated and non-treated mothers are followed over the same period. Finally, we estimate this new specification on a sample of mothers with at least two children who had their first one during the reform year. As we do that, we exclude from our analysis all the mothers for whom by construction we cannot observe any variation in birth spacing.

Table A8 in Appendix reports the estimates of this analysis. It first shows, as previously stated, no statistically significant effect for mothers above 30. On the other hand, the table shows that mothers, aged less than 30 years old and who had their first child with a father eligible for paternity leave in 2002, spent on average 35 fewer days on disability and received 1198 euros less in disability benefits after the birth of their second child. Interestingly, those effects are close in magnitude to the ones measured above for the whole length of the period following the birth of the first child in 2002. In other words, the effect that we observe in Table A7 for first-time mothers in the reform year and over a period of 12 years seems to match the effect we found here after the birth of the second child. This leads us to conclude that most of the overall reduction in disability occurred after the birth of the second child and that the decrease is driven by mothers who delayed the birth of their second child.

The rationale behind this argument is that increased birth spacing could have improved the health of the mothers, as well as their labor market attachment. Both of these improvements would have long-term consequences on disability. Regarding the first aspect, there is ample evidence from the medical literature that short birth spacing is detrimental to a woman's health. A recent review of 58 observational studies has shown that short intervals between pregnancies

[46]This new setting is illustrated in Figure A11.

were indeed associated with several adverse health conditions (Conde-Agudelo, Rosas-Bermudez, Castaño, & Norton, 2012). When it comes to labor market attachment, however, evidence is rather limited and inconclusive. An empirical study on Sweden by Karimi (2014), using miscarriages between the first and second births as an instrument, finds that longer birth intervals have positive long-run effects on income and wage rates. On the other hand, Troske and Voicu (2013), using data for the United-States, show that increasing the time between the first and second childbirth worsens labor market outcomes for mothers by reducing their probability of working full-time.

In conclusion, we provide evidence that increased birth spacing, exogenously induced by the introduction of a paternity leave in 2002, might have lowered the time that women spent on disability insurance in the long run. Of course, the association between the timing of the second birth and disability prevalence is correlational. However, the timing of the two effects match perfectly. In addition, the heterogeneity analyses have shown that the sub-populations driving the results are the same, that is young mothers who had a first child during the reform year. We also ruled out the potential mechanical effects by looking only at disability spells following the birth of the second child. Therefore, we conclude that the increase of time between births is the most likely candidate mechanism for the long-term reduction in disability observed after the introduction of paternity leave.

5 Conclusion

This paper examines how parenthood and parental gender impacts the probability of experiencing work disability at a young age. While previous work has highlighted the existence of "child penalties" related to women's earnings, we document another child penalty related to work disability that may prove to be equally important for a woman's career in the long run. Notably, the provision of paternity leave softens this "other child penalty" especially for first-time mothers under the age of 30.

Our study proceeds in the following direction. First, we use an event study approach to provide empirical evidence demonstrating that the incidence rate of work disability for women and men only begins to diverge after the birth of their first child. This gender gap in disability culminates over time so much that even eight years after childbirth women are 40% more likely to experience a disability that prevents them from working at their full capactiy. We also demonstrate that the impact of children on maternal health increases with the size of the family, with a gender gap in the probability to suffer disability that raises to 2.3 percentage points in families with three children. We believe that these results provide significant new insights into the career trajectories of mothers and the specific role that gaps due to poor health and disability play on their labor market attachment.

Drawing on this result, we next examine how the provision of paternity leave could moderate this so called “other” child penalty. We argue that if family arrangements related to childcare and a mother’s tendency to fall into disability are linked, then paternity leave provisions, which have been found to increase father’s involvement in child raising, could have a positive impact on maternal health. We exploit a discontinuity in Belgian legislation, which offered paternity leave only to fathers of children born after July 1st 2002, to evaluate the causal effect of the policy on the prevalence of work disability among mothers. Following a regression discontinuity difference-in-differences (RD-DiD) design, we find that mothers who gave birth to a child immediately after the reform spent on average 21% fewer days on disability over a period of 12 years. This result seems to be largely driven by younger women who had their first child during the reform year. Lastly, and with regards to the specific causes of maternal disability, our results show that mothers with musculoskeletal disorders spent 50% fewer days on disability insurance in the long term.

In conclusion, we provide suggestive evidence that an increase in birth spacing, induced by the paternity leave reform, could have played a large role in the reduction of the time that mothers spent on disability. We demonstrate that results connected to maternal disability and family planning are driven by the same sub-population of younger mothers who decided to delay the birth of their second child. We also provide evidence that both results exhibit similar time dynamics. This leads us to conclude that the timing of births for multiple-children families is key to reducing the problem of work disability of mothers at young ages.

Recent discussions at the European Union level indicate that our findings could provide useful insights in the context of the work-life balance directive, which was adopted by the European Council on June 13, 2019 and should be implemented in all members states within three years. The directive introduces a paternity leave of 10 days for fathers, which corresponds exactly to the laws currently in place in Belgium, making it a particularly interesting case for research. Our findings show that paternity leave policies might favor a convergence in gender inequalities, while reducing public spending on disability insurance programs. Those results are equally important for countries outside Europe, especially the United-States that has not yet adopted a nation-wide paid leave policy.

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Table 1: Descriptive Statistics and Balancing Test

		Sample statistics		**RD-DiD**		
		Mean	SD	Coeff.	SE	Obs.
Household†	Live in flanders (0/1)	0.65	(0.48)	0.00	(0.02)	101735
	Size (#)	3.76	(0.96)	-0.01	(0.03)	101735
	Children (#)	1.72	(0.84)	0.01	(0.03)	101735
	First child (0/1)	0.48	(0.50)	-0.02	(0.02)	101735
Mother	Age†	30.22	(4.14)	-0.11	(0.12)	101735
	Salaried employment (0/1)	0.90	(0.30)	0.02	(0.01)	101735
	Blue collar (0/1)	0.16	(0.37)	-0.02	(0.02)	84993
	White collar (0/1)	0.73	(0.44)	0.02	(0.02)	84993
	Civil servant (0/1)	0.11	(0.31)	0.00	(0.01)	84993
	Self-employed (0/1)	0.09	(0.28)	-0.01	(0.01)	101735
	Daily wage (euro)‡	78.90	(51.03)	0.45	(2.32)	51840
Father	Age†	32.50	(4.89)	-0.05	(0.10)	99502
	Salaried employment (0/1)	0.83	(0.38)	-0.01	(0.01)	99502
	Blue collar (0/1)	0.39	(0.49)	-0.01	(0.01)	77946
	White collar (0/1)	0.52	(0.50)	0.01	(0.01)	77946
	Civil servant (0/1)	0.09	(0.29)	0.00	(0.01)	77946
	Self-employed (0/1)	0.16	(0.36)	0.00	(0.00)	99502
	Daily wage (euro)‡	96.20	(56.12)	-0.34	(1.54)	50765

Notes: Columns 1-2 report means and standard deviations. Columns 3-5 report results from RD-DiD regressions based on equation (3) and using the reform (2002) and non-reform years (2003 -2004). All samples include fathers and mothers who were employed at the time of birth. Standard errors are clustered at birth month level. † Outcomes measured on Dec. 31 of each year. ‡ Outcomes measured the quarter before birth; sample limited to 3 months window. Since the data start in 2002, we cannot observe the outcomes for those who had a child between Jan. and March 2002.

Table 2: Effects of Paternity Leave Reform on Maternal Disability 12 years after Reference Child's Birth

	All mothers		**First-time mothers**		**Experienced mothers**	
	Coeff/SE	Mean	Coeff/SE	Mean	Coeff/SE	Mean
Panel A - Cumulative days on DI	-22.3 **	104.6	-38.6 ***	96.8	-7.4	111.7
	(8.9)		(13.4)		(13.3)	
Short-term (less than 12 months)	-6.3 **	55.4	-9.7 **	53.0	-3.1	57.5
	(3.0)		(4.7)		(3.9)	
Long-term (more than 12 months)	-16.1 **	49.2	-28.9 ***	43.8	-4.4	54.2
	(7.0)		(10.5)		(10.4)	
Panel B - Cumulative DI benefits	-712 **	4049	-1322 ***	3806	-159	4270
	(302)		(431)		(453)	
Short-term (less than 12 months)	-157	2194	-298 *	2146	-29	2238
	(119)		(162)		(150)	
Long-term (more than 12 months)	-555 **	1855	-1025 ***	1660	-130	2032
	(227)		(360)		(349)	
Panel C - Ever on DI	0.005	0.399	-0.005	0.407	0.014	0.393
	(0.012)		(0.014)		(0.017)	
Short-term (less than 12 months)	0.007	0.398	-0.004	0.405	0.016	0.391
	(0.012)		(0.014)		(0.017)	
Long-term (more than 12 months)	-0.009	0.061	-0.021 **	0.056	0.001	0.066
	(0.006)		(0.009)		(0.009)	
Number of observations	101,735		48,505		53,230	

Notes: This table reports RD-DiD estimates based on equation (3) and using the reform (2002) and non-reform years (2003 -2004). Outcomes in Panels A and B capture the cumulative effects over the 12-year period for the number of days and benefits, respectively. The variables labeled "Ever on DI" in Panel C are dummies for the probability to have entered disability insurance at least once over the 12-year period. Regressions control for mothers' age, number of children, as well as region of living, at the moment of the birth of the reference child. The sample includes mothers who were employed at the time of birth. Standard errors (reported in parentheses) are clustered at birth month level. Sample means are reported in the second column.

Table 3: Effects of Paternity Leave Reform on Maternal Disability (long-term only) - 11 years after Reference Child's Birth
(Heterogeneous effects by type of disease)

	All mothers	
	Coeff/SE	Mean
All conditions	-12.4 **	39.8
	(5.9)	
Mental disorders	-0.2	14.8
	(3.9)	
Diseases of musculoskeletal system and connective tissue	-5.7 **	9.6
	(2.7)	
Other	-6.4	15.4
	(5.4)	
Number of observations	101,735	

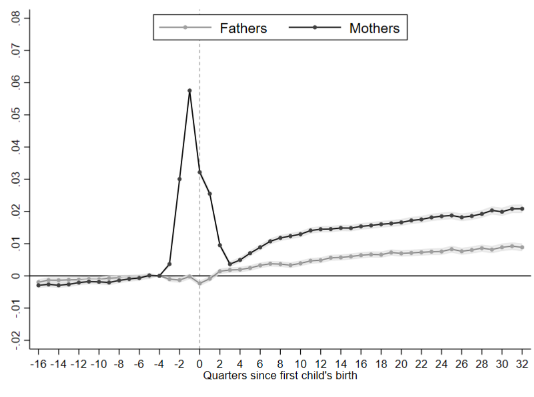
Notes: This table reports RD-DiD estimates based on equation (3) and using the reform (2002) and non-reform years (2003 -2004). Outcomes are for the long-term disability program only. Regressions control for mothers' age, number of children, as well as region of living, at the moment of the birth of the reference child. The sample includes mothers who were employed at the time of birth. Standard errors (reported in parentheses) are clustered at birth month level. Sample means are reported in the second column. Unlike the previous tables, we restrict the analysis to 11 years after childbirth because there was a change in the ICD classification of diseases in 2016 without any possibility to convert the data of that year to the previous classification system.

Table 4: Effects of Paternity Leave on Mothers' Subsequent Fertility

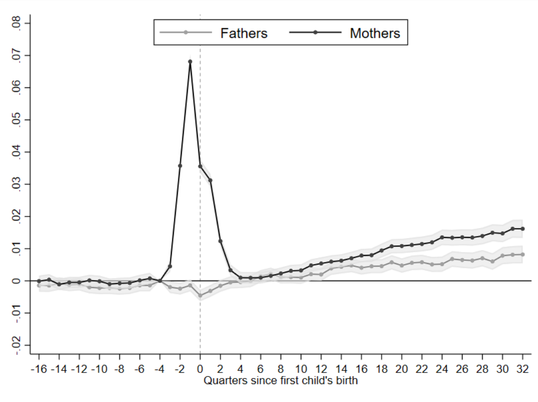
	First-time mothers < 30		**First-time mothers ≥ 30**	
	Coeff/SE	Mean	Coeff/SE	Mean
Other child				
After 1 year	-0.016	0,080	-0.021	0,077
	(0.010)		(0.013)	
After 2 years	-0.048 ***	0,360	-0.013	0,309
	(0.016)		(0.021)	
After 3 years	-0.055 **	0,584	-0.011	0,472
	(0.021)		(0.025)	
After 6 years	-0.018	0,775	-0.015	0,607
	(0.017)		(0.029)	
After 12 years	-0.008	0,831	0.003	0,641
	(0.016)		(0.027)	
Nb. children 12 years	-0.056	2,1	0.018	1,8
	(0.037)		(0.056)	
Number of observations	28.449		18.108	

Notes: This table reports RD-DiD estimates based on equation (3) and using the reform (2002) and non-reform years (2003 -2004). Regressions control for mothers' age, as well as region of living, at the moment of the birth of the reference child. The sample includes mothers who had a first child between 2002 and 2004 and were employed at the time of the birth of the reference child. The dependent variable "other child" is an indicator for the mother having another child within the following years after the reference child's birth. Standard errors (reported in parentheses) are clustered at birth month level. Sample means are reported in the second column.

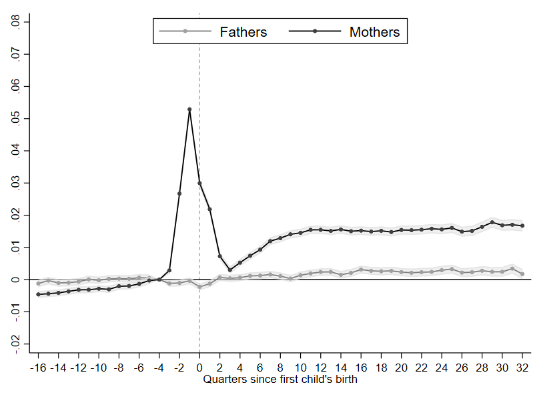
Panel A: All Parents



Panel B: One-child Parents



Panel C: Two-child Parents



Panel D: Three-child Parents

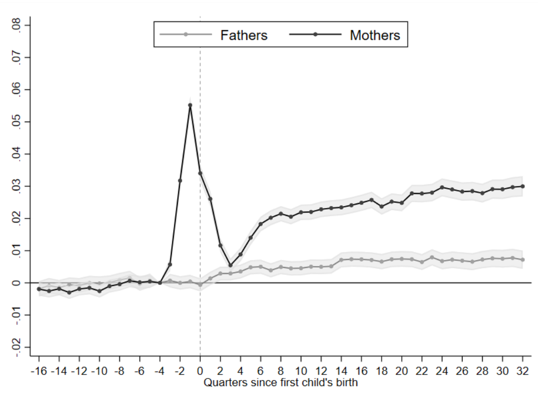


Figure 1: Impact of Children on Disability Receipt
(relative to event time -4)

Notes: The figures show event time coefficients for the probability to be on disability insurance (for both the short-term and long-term programs) relative to the 4^{th} quarter before the first child's birth, estimated from equation (1) for men and women separately. The sample includes all parents who had a first child between 2003 and 2013. For Panels B, C and D, we split the sample by the parents' total number of children as of 2016 (1,2 or 3 children). The shaded 95% confidence intervals are based on robust standard errors.

Panel A: Short-term Disability　　Panel B: Long-term Disability

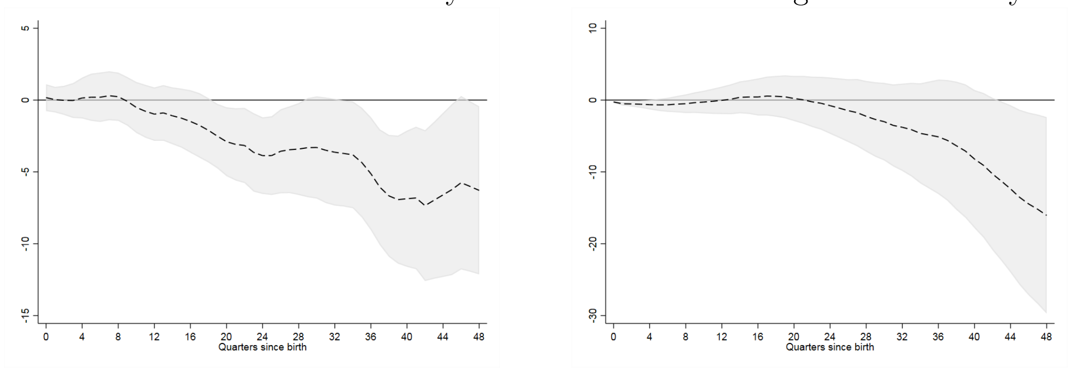


Figure 2: Cumulative Effects of Paternity Leave Reform on Mothers' Disability Days

Notes: The figures show RD-DiD estimates from 48 regressions based on equation (3). The sample includes mothers who were employed at the time of birth. The short-term program (Panel A) includes individuals who have spent less than 12 months on DI. The long-term program (Panel B) includes individuals who have spent more than 12 months on DI. The shaded 95% confidence intervals are based on clustered standard errors at birth month level.

Panel A: First-time Mothers   Panel B: Experienced Mothers

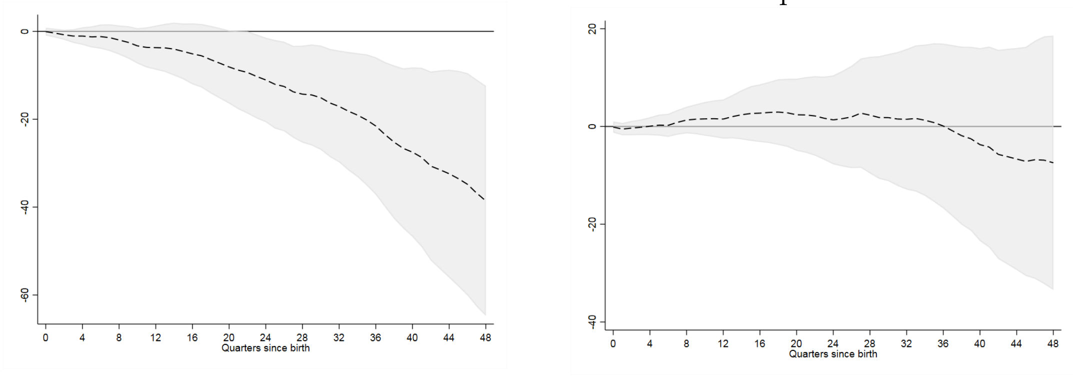
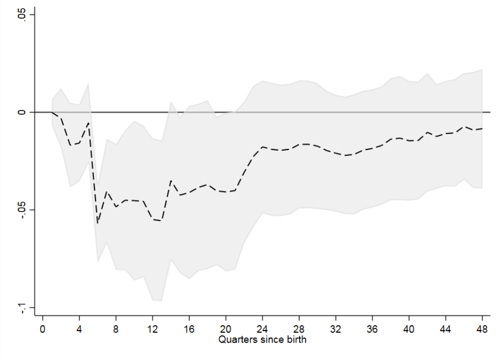


Figure 3: Cumulative Effects of Paternity Leave Reform on Mothers' Disability Days (Heterogeneous effects by birth order of reference child)

Notes: The figures show RD-DiD estimates from 48 regressions based on equation (3). The sample includes mothers who were employed at the time of birth. Results combine effects for both the short- and long-term programs. The shaded 95% confidence intervals are based on clustered standard errors at birth month level.

Panel A: Mothers $<$ 30 years old at the Birth of the Reference Child



Panel B: Mothers $\geq$ 30 years old at the Birth of the Reference Child

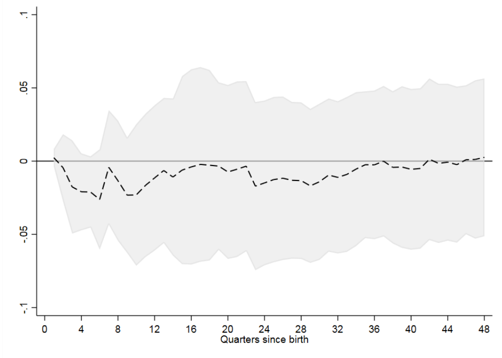


Figure 4: Causal Effects of the Paternity Leave on Mothers' Probability to have a Second Child

Notes: The figures show RD-DiD estimates from 48 regressions based on equation (3). The sample includes mothers who had a first child during the reform year and were employed at the time of birth. The shaded 95% confidence intervals are based on clustered standard errors at birth month level.

Appendix

Appendix - Table A1: Time use survey - Belgium 2013

	Men	Women	Diff.
Paid work	5:01	3:57	- 1:04
Household work	1:54	2:58	+ 1:04
Childcare and raising children	0:33	1:05	+ 0:32
Personal care	2:15	2:24	+ 0:09
Sleep and rest	8:17	8:29	+ 0:12
Education	0:06	0:06	+ 0:00
Social participation	1:14	1:10	- 0:04
Free time	3:13	2:23	- 0:50
Transportation	1:25	1:24	- 0:01
Other	0:03	0:05	+ 0:02

Notes: Household with both parents working and children.

Appendix - Table A2: Main features of the Belgian parental leave system

	Maternity leave	**Paternity leave**	**Parental leave**
Date introduction	1971	July 2002	1997
Legal texts	"Loi sur le travail", March 16, 1971	"Loi relative a la conciliation entre l'emploi et la qualité de vie", August 10, 2001 "Loi modifiant, en ce qui concerne les coparents, la législation afférente au congé de paternité", April 13, 2011 "Loi-programme du 22 décembre 2008", December 22, 2008	"Arreté royal relatif a l'introduction d'un droit au congé parental dans le cadre d'une interruption de la carrišre professionnelle", October, 29 1997
Duration	Max. 15 weeks (min. 1 before planned delivery + 9 after childbirth)	Max. 2 weeks (initially to be taken within 1 month after childbirth, extended to 4 months in April 2009)	Max. 4 months (per parent per child)
Conditions	Only women who worked min. 120 days in last 6 months	Only fathers (and co-parent in same sex couples since 2011) with salaried contract	Leave may be taken full-time, half-time over 8 months or for one day a week (one-fifth-time) over 20 months Leave may be taken up to the child's 12th birthday Both parents can take leave at the same time
Replacement rate	82% gross salary (first 30 days) 75% remaining days (capped)	First 3 days fully compensated Remain. 7 days 82% gross salary	Flat rate 802 euros per month

Appendix - Table A3: Bunching in number of births at the threshold

Window	7 days	14 days	21 days	28 days	35 days	42 days
Log n. of births	0.030	0.029	0.032	0.013	-0.013	0.012
	(0.028)	(0.051)	(0.041)	(0.031)	(0.028)	(0.036)
Linear trend	N	Y	Y	Y	Y	Y
Day of the week	Y	Y	Y	Y	Y	Y
Number of observations	14	28	42	56	70	84

*Notes: This table reports RDD estimates from regressions of the form of equation (2). The outcome variable is the log daily number of births. The reported coefficients are from a binary indicator for birthdates on or after July 1st, 2002. The sample includes all days in the specified window around the date of the introduction of the paternity leave. In all but the first column, we control for a linear trend in date of birth (i.e. the running variable, centered at 0 in July 1st, 2002), interacted with the binary indicator. Robust standard errors are reported in parentheses. Data source: Belgian statistical office - StatBel. *** p<0.01, ** p<0.05, * p<0.1.*

Appendix - Table A4: Effects of paternity leave reform on paternal disability - 12 years after the reference child's birth

	Coeff/SE	Mean
Panel A - Cumulative days on DI	-2.6	63.8
	(5.968)	
Short-term (less than 12 months)	-4.3	36.9
	(2.749)	
Long-term (more than 12 months)	1.7	26.9
	(4.175)	
Panel B - Cumulative DI benefits	-98.2	2929
	(250.7)	
Short-term (less than 12 months)	-176.4	1844
	(142.7)	
Long-term (more than 12 months)	78.2	1085
	(158.4)	
Panel C - Ever on DI	-0.029 ***	0.312
	(0.0)	
Short-term (less than 12 months)	-0.028 ***	0.311
	(0.0)	
Long-term (more than 12 months)	-0.001	0.034
	(0.0)	
Number of observations	99,502	

*Notes: This table reports RD-DiD estimates based on equation (3) and using the reform (2002) and non-reform years (2003 -2004). Outcomes in Panels A and B capture the cumulative effects over the 12-year period for the number of days and benefits, respectively. The variables labeled "Ever on DI" in Panel C are dummies for the probability to have entered disability insurance at least once over the 12-year period. Regressions control for fathers' age, number of children, as well as region of living, at the moment of the birth of the reference child. The sample includes fathers who were employed at the time of birth. Standard errors (reported in parentheses) are clustered at birth month level. Sample means are reported in the second column. *** p<0.01, ** p<0.05, * p<0.1.*

Appendix - Table A5: Effects of paternity leave reform on maternal disability - 12 years after reference child's birth (varying bandwidth)

	6	**5**	**4**	**3**	**2**	**1**	**Donut-hole**
	Coeff/SE	Coeff/SE	Coeff/SE	Coeff/SE	Coeff/SE	Coeff/SE	Coeff/SE
Panel A - Cumulative days on DI	-22.3 **	-19.5 **	-20.7 **	-28.1 ***	-31.2 ***	-20.5 ***	-19.2
	(8.9)	(7.9)	(7.4)	(7.5)	(9.2)	(3.4)	(11.9)
Short-term (less than 12 months)	-6.3 **	-4.5 **	-3.0	-7.3 **	-3.7 *	-3.9 ***	-8.0 *
	(3.0)	(2.1)	(2.1)	(2.7)	(1.9)	(0.2)	(3.9)
Long-term (more than 12 months)	-16.1 **	-15.0 **	-17.7 **	-20.7 ***	-27.4 ***	-16.6 ***	-11.2
	(7.0)	(6.9)	(7.0)	(6.9)	(8.5)	(3.2)	(9.2)
Panel B - Cumulative DI benefits	-712 **	-588 **	-500 *	-833 ***	-925 **	-626 ***	-677
	(302)	(246)	(280)	(282)	(371)	(148)	(425)
Short-term (less than 12 months)	-157	-92	3	-188	-20	-68	-241
	(119)	(86)	(104)	(113)	(118)	(40)	(165)
Long-term (more than 12 months)	-555 **	-496 **	-504 *	-645 **	-905 ***	-559 ***	-436
	(227)	(214)	(248)	(233)	(279)	(109)	(301)
Panel C - Ever on DI	0.005	0.022 **	0.022	0.016	-0.012	0.003	0.003
	(0.012)	(0.010)	(0.013)	(0.014)	(0.013)	(0.006)	(0.017)
Short-term (less than 12 months)	0.007	0.024 **	0.024 *	0.018	-0.008	0.005	0.004
	(0.012)	(0.010)	(0.013)	(0.014)	(0.013)	(0.006)	(0.017)
Long-term (more than 12 months)	-0.009	-0.011 **	-0.011 **	-0.011 *	-0.004	-0.008 ***	-0.011
	(0.006)	(0.004)	(0.004)	(0.005)	(0.004)	(0.000)	(0.010)
Number of observations	101,735	84,987	68,902	51,840	34,784	17,424	84,311

*Notes: This table reports RD-DiD estimates based on equation (3) and using the reform (2002) and non-reform years (2003 -2004). Outcomes in Panels A and B capture the cumulative effects over the 12-year period for the number of days and benefits, respectively. The variables labeled "Ever on DI" in Panel C are dummies for the probability to have entered disability insurance at least once over the 12-year period. Regressions control for mothers' age, number of children, as well as region of living, at the moment of the birth of the reference child. The sample includes mothers who were employed at the time of birth. Standard errors (reported in parentheses) are clustered at birth month level. Sample means are reported in the second column. *** p<0.01, ** p<0.05, * p<0.1.*

Appendix - Table A6: Effects of paternity leave reform on maternal disability - 12 years after reference child's birth (varying polynomial order)

	Linear Coeff/SE	**Quadratic** Coeff/SE	**Cubic** Coeff/SE
Panel A - Cumulative days on DI	-22.3 **	-20.7 *	-51.5 **
	(8.9)	(11.5)	(19.6)
Short-term (less than 12 months)	-6.3 **	-1.6	-11.1
	(3.0)	(3.2)	(7.2)
Long-term (more than 12 months)	-16.1 **	-19.2 *	-40.4 **
	(7.0)	(10.6)	(16.9)
Panel B - Cumulative DI benefits	-712 **	-466	-1541 *
	(302)	(424)	(782)
Short-term (less than 12 months)	-157	54	-267
	(119)	(147)	(331)
Long-term (more than 12 months)	-555 **	-520	-1275 **
	(227)	(358)	(612)
Panel C - Ever on DI	0.005	0.041 **	-0.057
	(0.012)	(0.019)	(0.035)
Short-term (less than 12 months)	0.007	0.045 **	-0.055
	(0.012)	(0.019)	(0.035)
Long-term (more than 12 months)	-0.009	-0.013 *	0.000
	(0.006)	(0.007)	(0.013)
Number of observations	101,735	101,735	101,735

*Notes: This table reports RD-DiD estimates based on equation (3) and using the reform (2002) and non-reform years (2003 -2004). Outcomes in Panels A and B capture the cumulative effects over the 12-year period for the number of days and benefits, respectively. The variables labeled "Ever on DI" in Panel C are dummies for the probability to have entered disability insurance at least once over the 12-year period. Regressions control for mothers' age, number of children, as well as region of living, at the moment of the birth of the reference child. The sample includes mothers who were employed at the time of birth. Standard errors (reported in parentheses) are clustered at birth month level. Sample means are reported in the second column. *** p<0.01, ** p<0.05, * p<0.1.*

Appendix - Table A7: Effects of paternity leave reform on maternal disability - 12 years after reference child's birth
(heterogeneous effects by age of mother at birth)

	First-time mothers under 30		**First-time mothers over 30**	
	Coeff/SE	Mean	Coeff/SE	Mean
Panel A - Cumulative days on DI	-44.5 **	95.2	-25.6	99.3
	(18.3)		(17.4)	
Short-term (less than 12 months)	-12.0 *	56.6	-4.2	47.4
	(5.9)		(6.2)	
Long-term (more than 12 months)	-32.5 **	38.6	-21.3	51.9
	(13.9)		(14.4)	
Panel B - Cumulative DI benefits	-1687 **	3654	-633	4046
	(644)		(539)	
Short-term (less than 12 months)	-412 *	2205	-65	2053
	(222)		(257)	
Long-term (more than 12 months)	-1274 **	1449	-568	1993
	(480)		(462)	
Panel C - Ever on DI	-0.021	0.437	0.026	0.359
	(0.022)		(0.027)	
Short-term (less than 12 months)	-0.019	0.436	0.027	0.357
	(0.022)		(0.027)	
Long-term (more than 12 months)	-0.023 *	0.054	-0.017	0.058
	(0.012)		(0.012)	
Number of observations	29,648		18,857	

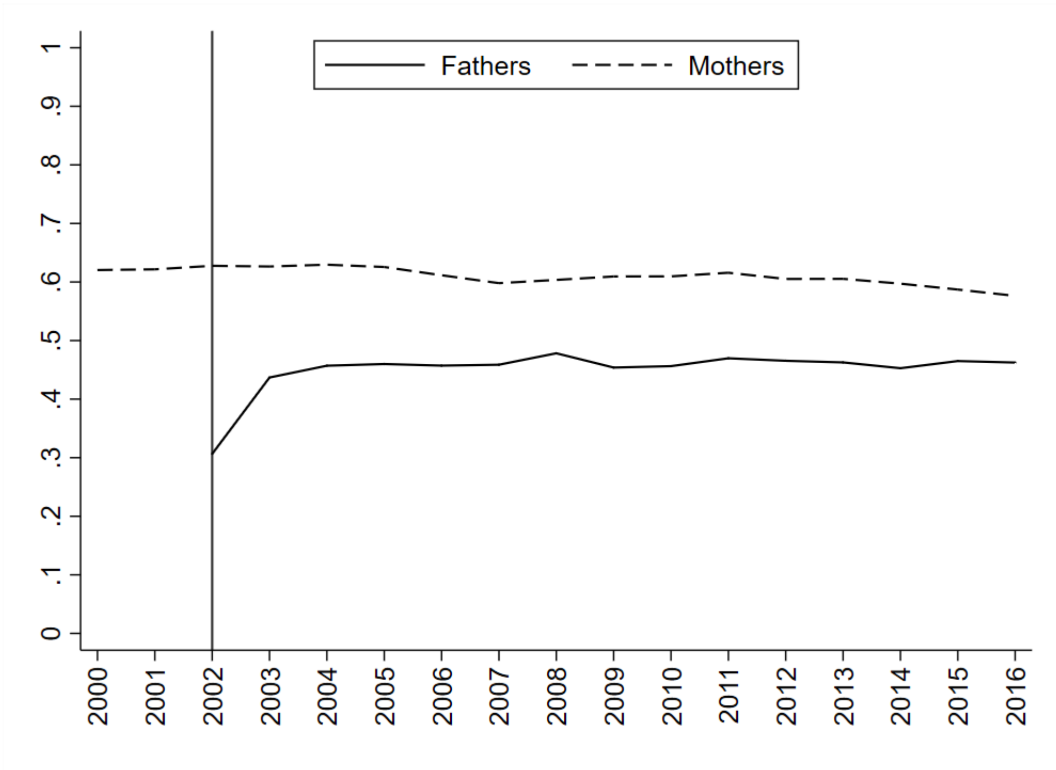
*Notes: This table reports RD-DiD estimates based on equation (3) and using the reform (2002) and non-reform years (2003 -2004). Outcomes in Panels A and B capture the cumulative effects over the 12-year period for the number of days and benefits, respectively. The variables labeled "Ever on DI" in Panel C are dummies for the probability to have entered disability insurance at least once over the 12-year period. Regressions control for mothers' age, number of children, as well as region of living, at the moment of the birth of the reference child. The sample includes mothers who had a first child during the reform year and were employed at the time of birth. Standard errors (reported in parentheses) are clustered at birth month level. Sample means are reported in the second column. *** p<0.01, ** p<0.05, * p<0.1.*

Appendix - Table A8: Effects of paternity leave reform on maternal disability - 8 years after the second child's birth

	First-time mothers under 30		First-time mothers over 30	
	Coeff/SE	Mean	Coeff/SE	Mean
Panel A - Cumulative days on DI	-34.8 **	55.9	4.5	49.7
	(16.9)		(13.7)	
Short-term (less than 12 months)	-10.7	35.6	-0.4	27.1
	(7.3)		(4.7)	
Long-term (more than 12 months)	-24.2 **	20.4	4.9	22.6
	(10.9)		(11.0)	
Panel B - Cumulative DI benefits	-1198 **	2095	508	2038
	(528)		(579)	
Short-term (less than 12 months)	-367	1368	99	1196
	(251)		(216)	
Long-term (more than 12 months)	-831 **	726	409	842
	(322)		(450)	
Panel C - Ever on DI	-0.010	0.329	0.049	0.254
	(0.021)		(0.038)	
Short-term (less than 12 months)	-0.011	0.328	0.046	0.251
	(0.020)		(0.039)	
Long-term (more than 12 months)	-0.020	0.035	-0.003	0.034
	(0.015)		(0.010)	
Number of observations	21,646		10,909	

*Notes: This table reports RD-DiD estimates based on equation (3) and using the reform (2002) and non-reform years (2003 -2004). Outcomes in Panels A and B capture the cumulative effects over the 8-year period for the number of days and benefits, respectively. The variables labeled "Ever on DI" in Panel C are dummies for the probability to have entered disability insurance at least once over the 8-year period. Regressions control for mothers' age, number of children, as well as region of living, at the moment of the birth of the reference child. The sample includes mothers who had a first child during the reform year, were employed at the time of the first birth, and had at least another child in the years following the reform. Standard errors (reported in parentheses) are clustered at birth month level. Sample means are reported in the second column. *** p<0.01, ** p<0.05, * p<0.1.*

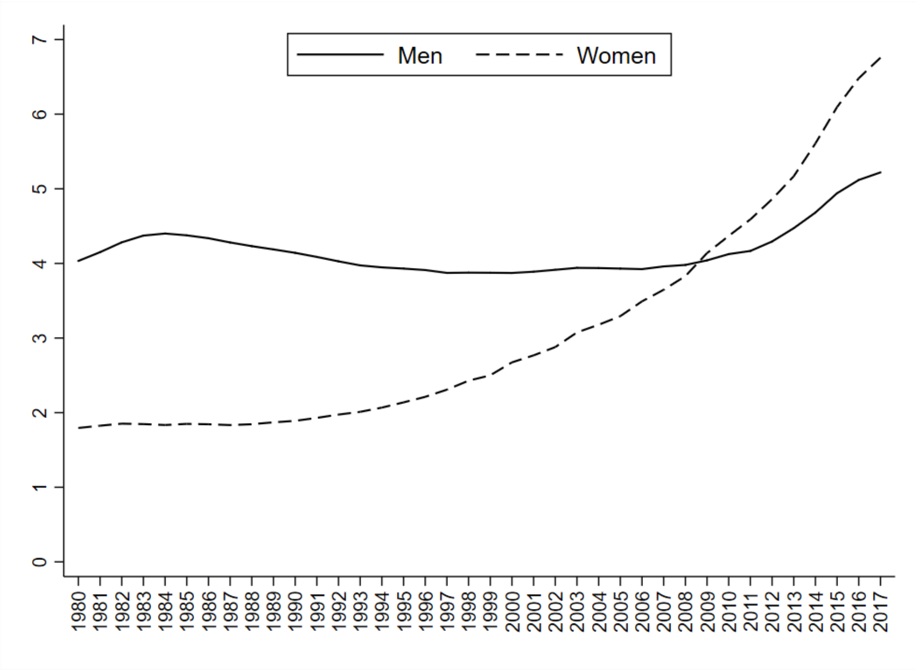
Appendix - Figure A1: Number of fathers/mothers taking paternity/maternity leave as a fraction of the annual number of births



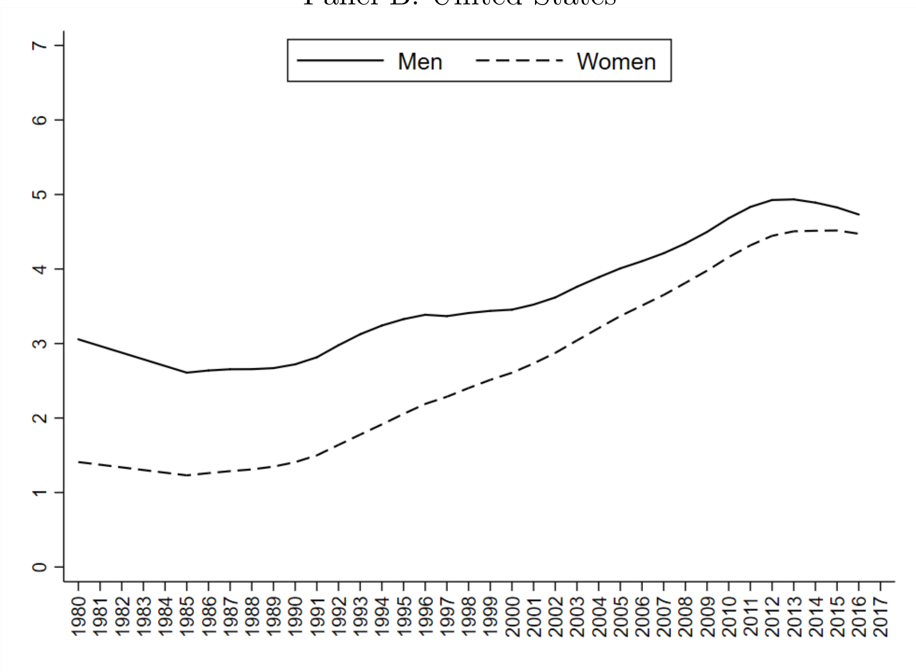
Notes: Statistics for maternity leave do not include civil servants and self-employed workers who benefit from a different system. It should also be noticed that women who have not worked at least 120 days during the last 6 months are not entitled to maternity leave. For all these reasons, only 60% of women on average are reported to have taken a maternity leave over the last two decades. Statistics for paternity leave do not include fathers who stop working for only 3 days or less since they need only to report to their employer. One should also keep in mind that those statistics do not account for civil servants, who benefit from a different system, as well as self-employed workers who were not entitled to paid paternity leave before 2019. For the year of the reform, we only consider births from July to December 2002. Data sources: National Institute for Health and Disability Insurance (leave-takers) and StatBel (births).

Appendix - Figure A2: Percent of working-age (20–64) population receiving (long-term) DI benefits

Panel A: Belgium



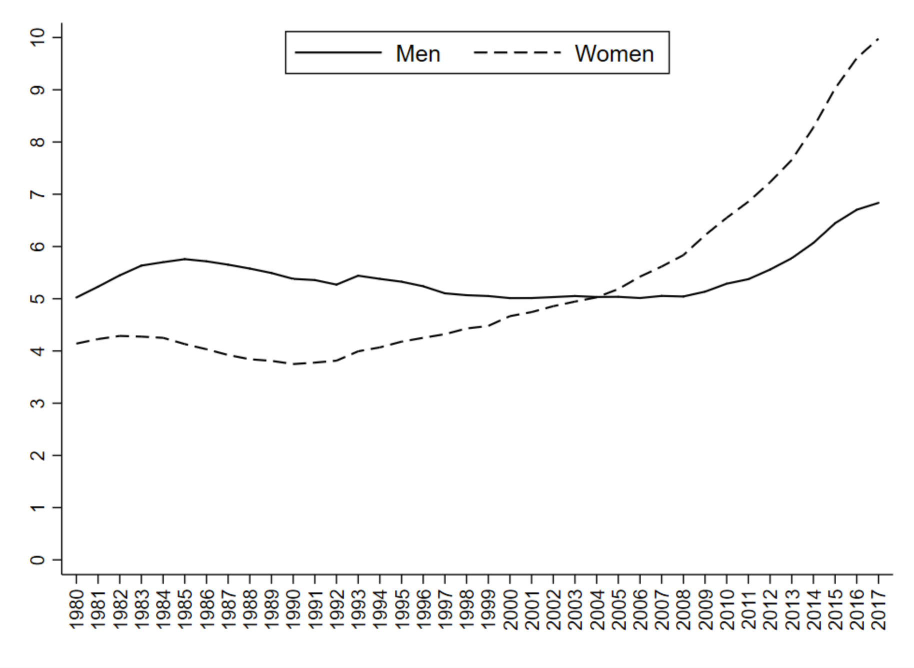
Panel B: United States



Data source for Belgium: National Institute for Health and Disability Insurance & OECD. Data source for the United States: Social Security Administration, 2017 Annual Statistical Supplement & OECD.

Appendix - Figure A3: Percent of insured workers receiving (long-term) DI benefits

Panel A: Belgium

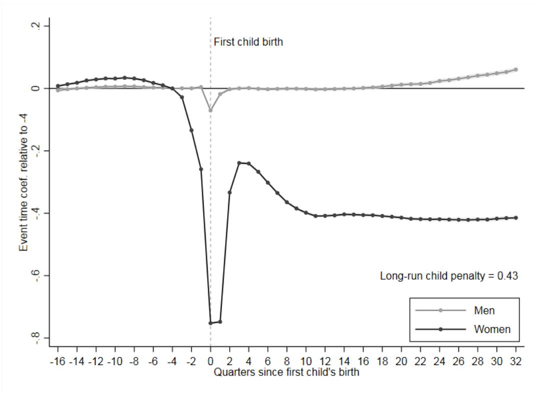


Panel B: United States

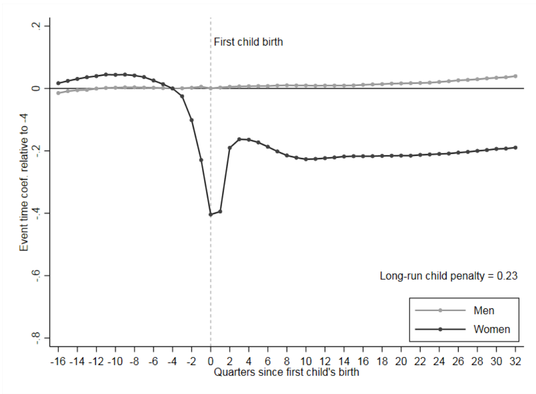
Data source for Belgium: National Institute for Health and Disability Insurance. Data source for the United States: Social Security Administration, 2017 Annual Statistical Supplement.

Appendix - Figure A4: Impacts of children

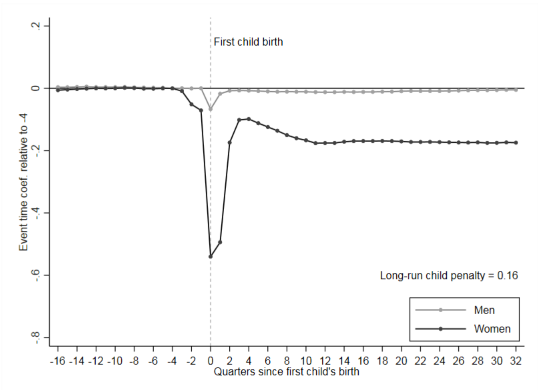
Panel A: Earnings



Panel B: Participation rates

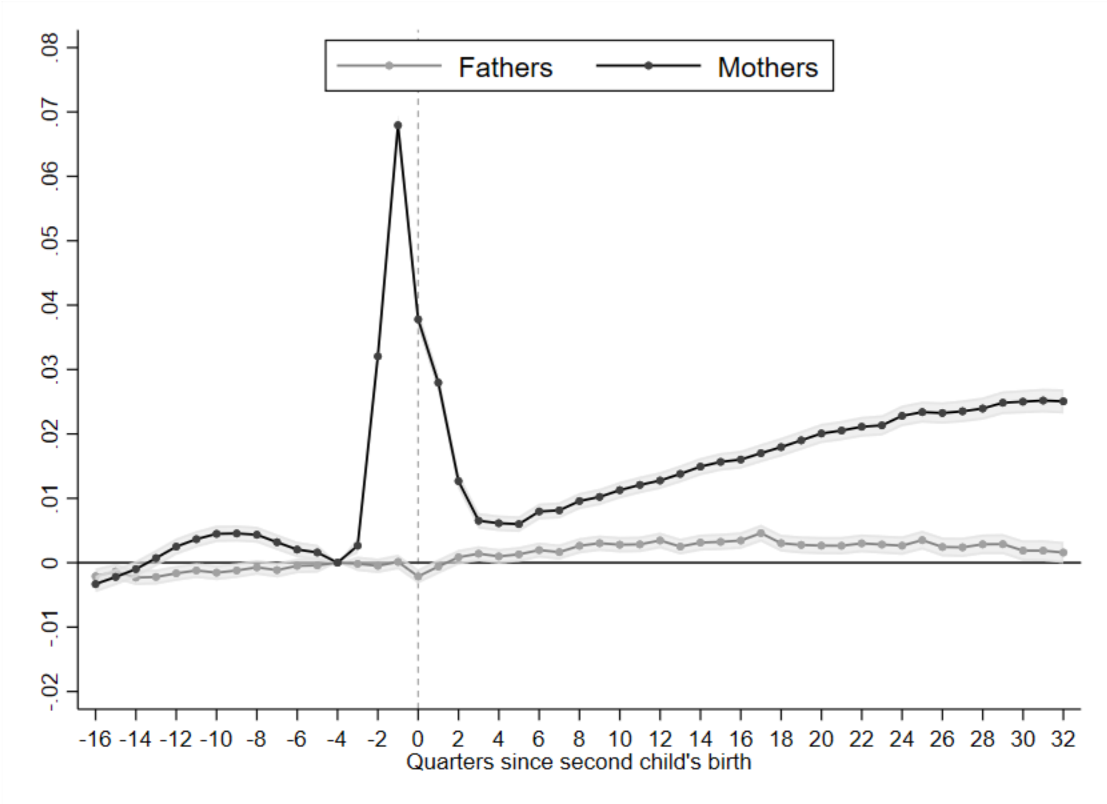


Panel C: Hours worked



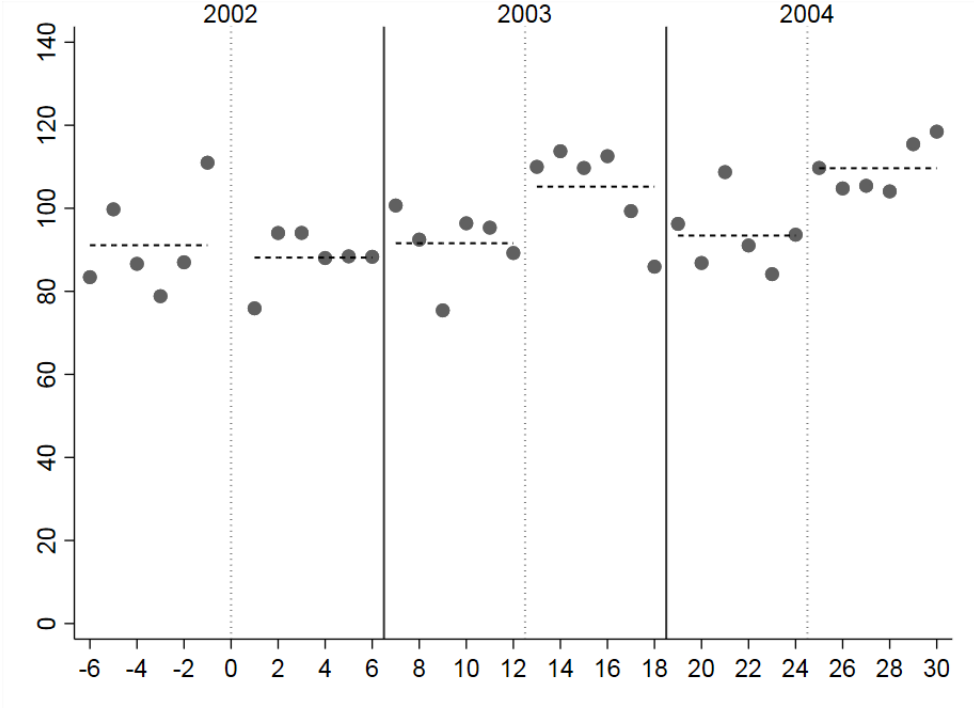
Notes: The figures show event time coefficients estimated from equation (1) relative to the 4th quarter before the first child's birth, for men and women separately. The coefficients are displayed as a percentage of the mean of the outcome measured at t-4. We report results for gross labor earnings (excluding taxes or transfers), participation rates and hours worked. The effects on earnings and participation are estimated unconditional on employment status, while the effects on hours worked are estimated conditional on participation. The long-run child penalty - the percentage by which women are falling behind men due to children - is defined as the average penalty from event time 12 to 32. The sample includes all parents who had a first child between 2002 and 2013. The shaded 95% confidence intervals are based on robust standard errors.

Appendix - Figure A5: Event study around second child's birth - Impact on disability receipt (relative to event time -4)



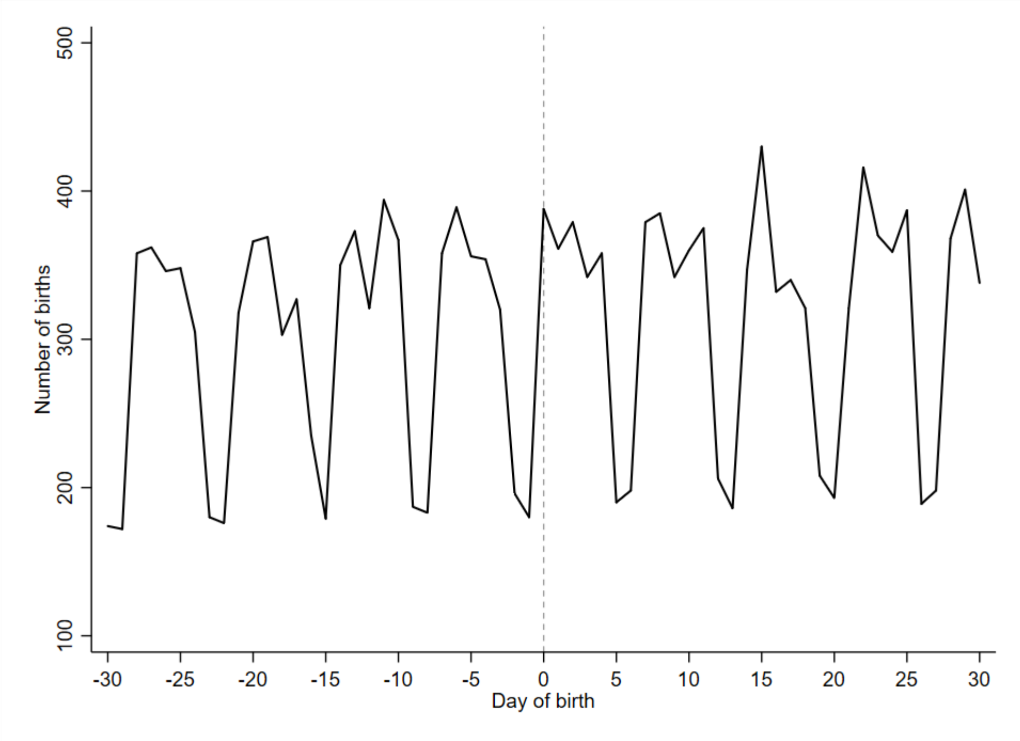
Notes: The figure shows event time coefficients for the probability to be on disability insurance (for both the short-term and long-term programs) relative to the 4^{th} quarter before the second child's birth, estimated from equation (1) for men and women separately. t=0 is now the quarter of birth of the second child. All of these statistics are estimated on a sample of parents who have had two children in total as of 2016. The shaded 95% confidence intervals are based on robust standard errors.

Appendix - Figure A6: Seasonality - Total number of disability days 12 years after the reference child's birth



Notes: Sample of mothers who had a first child in 2002-2004. The horizontal dashed lines represent the average within a given semester. Time (horizontal axis) is indexed on the introduction of the paternity leave on the 1st of July 2002.

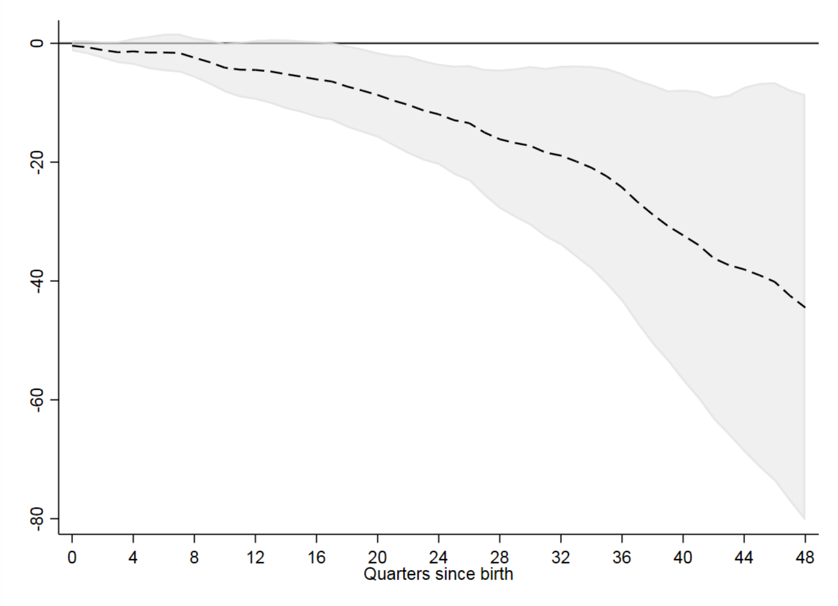
Appendix - Figure A7: Daily number of births



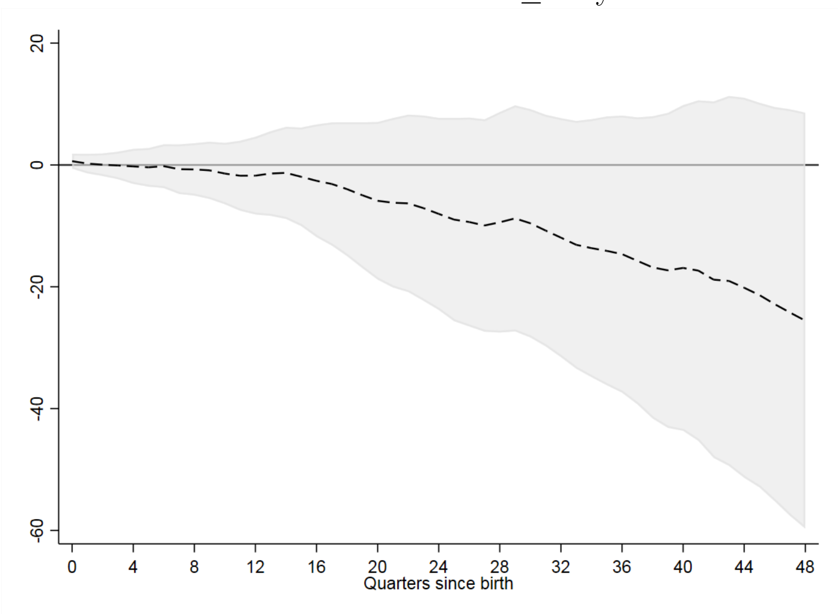
Notes: Daily number of births around the introduction of paternity leave. The day of birth is normalized to 0 for July 1st, 2002. Data source: Belgian statistical office - StatBel.

Appendix - Figure A8: Cumulative effects of paternity leave reform on mothers' disability days
(heterogeneous effects by age of mother at birth of the reference child)

Panel A: First-time mothers < 30 years old



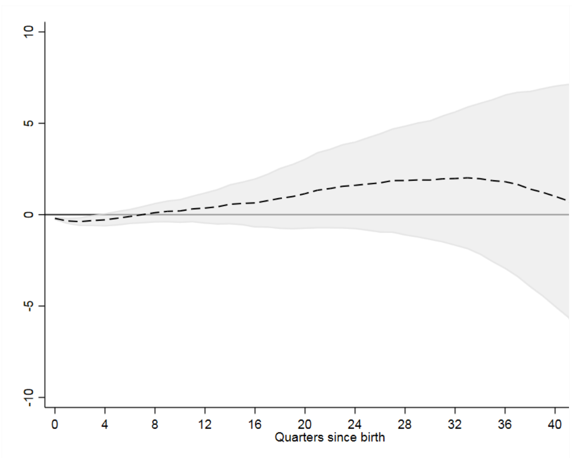
Panel B: First-time mothers ≥ 30 years old



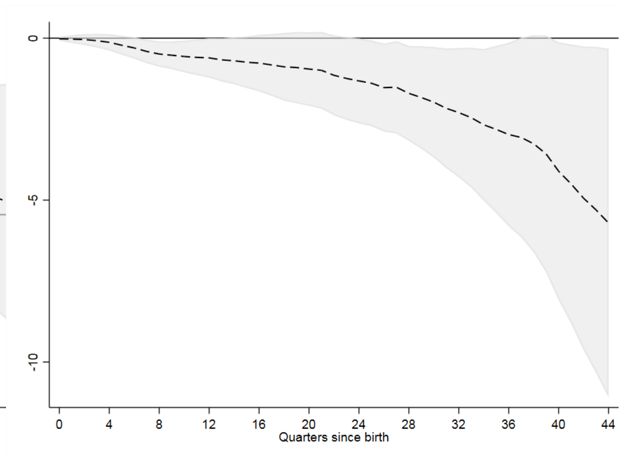
Notes: The figures show RD-DiD estimates from equation (3). All of these statistics are estimated on a sample of mothers who had a first child between 2002 and 2004 and were employed at the time of the birth of the reference child. The shaded 95% confidence intervals are based on clustered standard errors at birth month level.

Appendix - Figure A9: Cumulative effects of paternity leave reform on mothers' long-term disability days
(heterogeneous effects by medical condition)

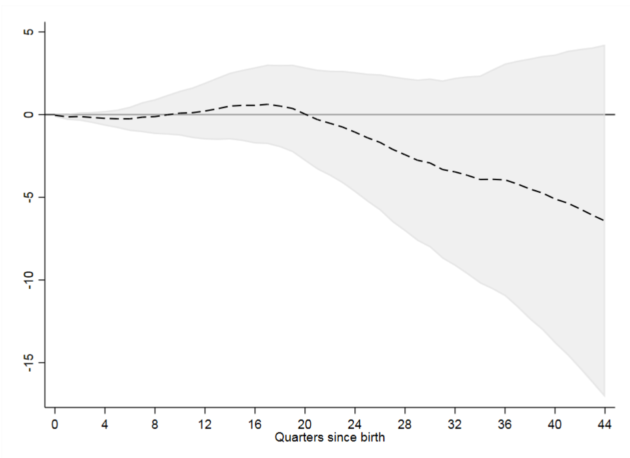
Panel A: Mental and behavioral disorders



Panel B: Musculoskeletal system or connective tissue diseases



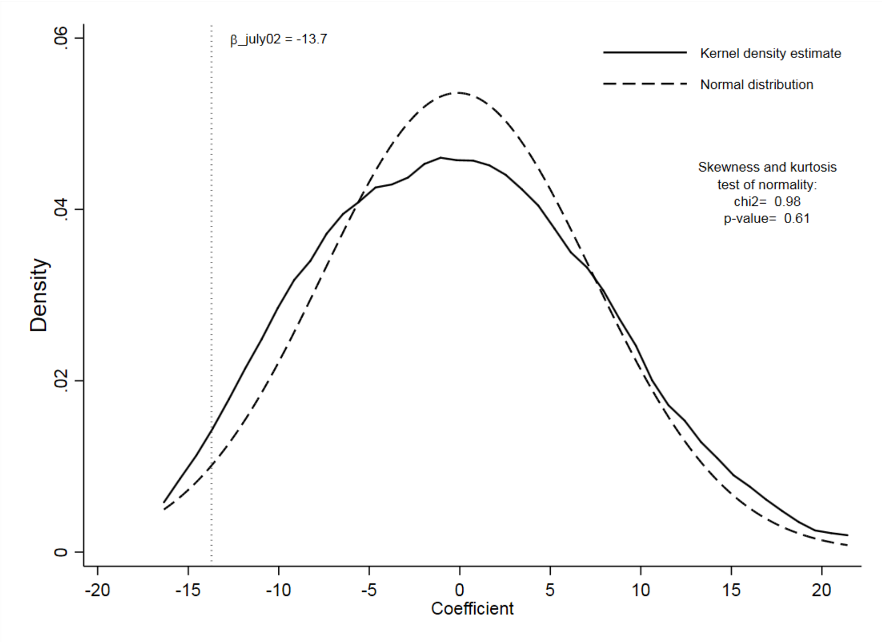
Panel C: Other diseases



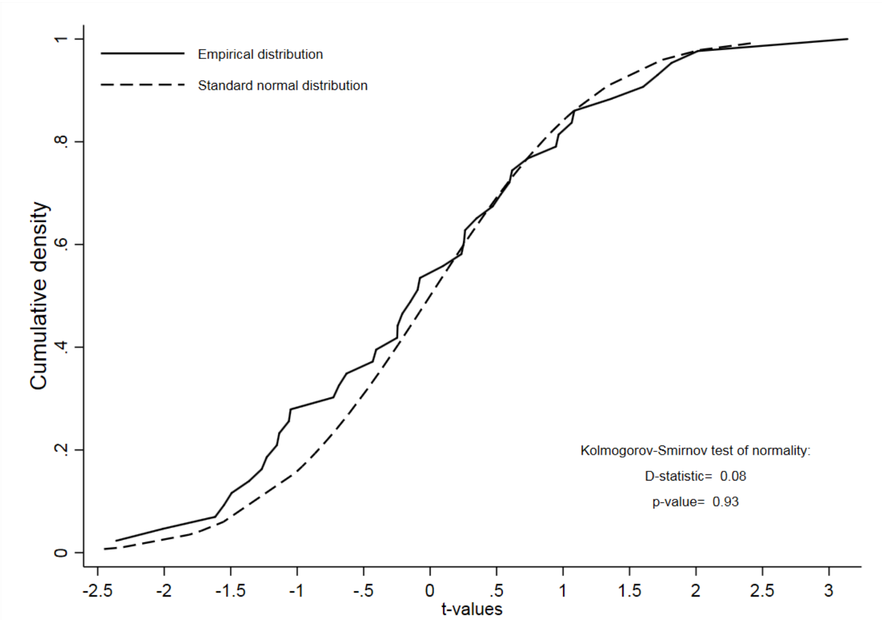
Notes: The figures show RD-DiD estimates from 48 regressions based on equation (3). Outcomes are for the long-term disability program only. All of these statistics are estimated on a sample of mothers who had a child between 2002 and 2004 and were employed at the time of the birth of the reference child. The shaded 95% confidence intervals are based on clustered standard errors at birth month level. Unlike the previous graphs, we restrict the analysis to 11 years after childbirth because there was a change in the ICD classification of diseases in 2016 without any possibility to convert the data of that year to the previous classification system.

Appendix - Figure A10: Placebo estimates for mothers' cumulative disability days after 10 years

Panel A: Placebo estimates



Panel B: t-values from placebo estimates



Notes: The figures show RD-DiD estimates from 43 regressions based on equation (3). All of these statistics are estimated on a sample of mothers who had a child between 2002-2006 and were employed at the time of the birth of the reference child.

Appendix - Figure A11: New follow-up period indexed on second child's birth