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Can Public Policies Break the Gender Mold? Evidence from Paternity Leave Reforms in Six Countries

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Can Public Policies Break the Gender Mold? Evidence from Paternity Leave Reforms in Six Countries

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Abstract

We investigate the impact of paternity leave policies on gender role attitudes in the next generation. We measure gender-stereotypical attitudes using an Implicit Association Test with 3,000 online respondents in six countries. Using an RD design, we observe a significant reduction (-0.21 SD) in gender-stereotypical attitudes among men born post-paternity leave implementation. This shift influences career choices, as men whose fathers were affected by the reform are more inclined to pursue counter-stereotypical jobs, particularly in high-skilled occupations like healthcare and education. Our findings highlight how paternity leave fosters egalitarian gender norms and affects the occupational choices of the next generation.

JEL Classification: J08, J13, J16, J18
Keywords : gender norms, paternity leave, female-dominated occupation, regression discontinuity

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A growing body of evidence demonstrates that the persistence of gender inequalities in labor market outcomes can be attributed to the strong specialization of men in market work and women in home production following the arrival of children (see e.g. Lundborg, Plug, & Rasmussen, 2017). This suggests that the normative roles of males-as-breadwinners and females-as-homemakers still very much drive career choices.

In response, public policies have been designed to favor the convergence of men's and women's roles, with the aim of lifting barriers to gender equality in economic outcomes. In recent decades, many policy makers have focused on paternity leave as a way to rebalance the traditional roles of parents in the household.[1] Recent evaluations show that paternity leave reforms have successfully increased the presence of fathers in the home after childbirth (Cools, Fiva, & Kirkebøen, 2015; Persson & Rossin-Slater, forthcoming), as well as their long-term involvement in childcare and household chores (Arnalds, Eydal, & Gíslason, 2013; Farré & González, 2019; Fontenay & Tojerow, 2024; Kotsadam & Finseraas, 2011; Pailhé, Solaz, & Tô, 2024; Patnaik, 2019; Tamm, 2019).

One important question that remains is the extent to which the greater involvement of fathers in child-rearing can promote more egalitarian attitudes in the next generation. And so, in this paper, we evaluate whether public policies, such as paternity leave, can promote counter-stereotypical attitudes that are transmitted from parents to children.

We use data on young adults exposed (or not) to a father who was eligible for paternity leave following reforms in six European countries with different cultural, social, and economic backgrounds: Belgium, Denmark, France, Iceland, Norway, and Sweden. We measure attitudes

[1] "Paternity leave" is defined as a policy that allows men to take time off from their job after the birth of a child, while subsidizing their earnings and allowing them to return to said job. We focus on parental leave policies that reserve time off for the father, thus not including leave entitlements that can be shared by both parents.

about gender roles using data from "Project Implicit,"[2] which provides publicly available Implicit Association Tests (IAT). In particular, we use data from the "Gender-Career" association test, which aims at measuring subconscious associations between women and family and between men and career. The main advantage of using IATs is avoiding the "social desirability bias" that arises when respondents answer survey questions in a manner that would be perceived more favorably by others (Carlana, Ferrara, & Pinotti, 2022). For this reason, previous research reveals that IATs are better at capturing a person's attitudes than survey questionnaires (Egloff & Schmukle, 2002).

Using a Regression Discontinuity Design, we compare young adults born right before and right after the reforms that first introduced paternity leave in six countries during the 1990's and early 2000's. We find that men exposed to eligible fathers display less gender-stereotypical attitudes, that is, they are less likely to associate women with home-making and men with career. The magnitude of the effect is meaningful since it corresponds to a reduction of 0.2 standard deviation across countries, where Sweden (-0.3 std.) and Belgium (-0.5 std.) stand out. Only men appear to be affected by the counter-stereotypical example that a father on paternity leave provides, while we do not observe any change in women's attitudes in any of the countries considered.

Building on this result, we examine how the change in gender role attitudes could translate into counter-stereotypical behaviors in the labor market. Using data from the European Union survey on Statistics on Income and Living Conditions (SILC), we find that treated male respondents (born right after the paternity leave reforms) living in Denmark, Norway, and

[2] Project implicit is a multi-university initiative founded by Dr. Tony Greenwald (University of Washington), Dr. Mahzarin Banaji (Harvard University) and Dr. Brian Nosek (University of Virginia). It is a non-profit organization providing a "virtual laboratory" for collecting data on implicit social cognition on the internet. Currently, 18 implicit association tests are hosted on the Project Implicit's website and have received IRB approval from the University of Virginia. For more information: https://www.projectimplicit.net/resources/about-the-iat/

Sweden, are more likely to work in a female-dominated occupation (by 4.5 percentage points), with the effect again most pronounced in Sweden, as well as in high-skilled occupations (e.g. health and teaching professionals). We argue that, by changing gender norms, paternity leave reforms can contribute to reducing gender segregation in the labor market. This seems particularly important since "occupational ghettos" have been found to be very persistent over time (Charles & Grusky, 2005), even in countries where women's participation in the labor market has reached levels comparable to that of men.

Finally, using survey data on weekly contribution to household chores, we provide suggestive evidence that paternity leave reforms have changed the division of labor in the long run. We posit that the long-term involvement of fathers in their home has helped to shape the next generation's gender role attitudes and contributed to their performance in the IAT. In other words, counter-stereotypical fathers seem to have acted as a role model for their sons.

We believe that this study fills an important gap since preference formation and social norms have been found to be an important contributor to the persistence of gender inequality across generations (Nollenberger, Rodríguez-Planas, & Sevilla, 2016). In addition, while a growing literature evaluates the effects of paternity leave reforms on parents, little is known about the extent to which the effects are transmitted to the next generation. To the best of our knowledge, only two studies, still unpublished, consider potential spillover effects of paternity leave from parents to children. Farré, Felfe, González, and Schneider (2021) find that at age 12, children whose fathers were eligible for paternity leave in Spain exhibit more egalitarian attitudes towards gender roles and engage more in counter-stereotypical behaviors at home. Mikkelsen and Peter (2022) reveal that the Swedish "daddy month" reform of 1995, which provided 30 days of earmarked parental leave for fathers, increased the probability that girls choose math-intensive programs in secondary education.

We contribute to this burgeoning literature on the intergenerational effects of paternity leave along three dimensions. First, we provide the first cross-country evidence of an intergenerational effect of paternity leave on gender norms. We believe that our diverse pool of countries, from Scandinavia to Western Europe, provides crucial external validity. Second, we are the first to consider effects on adults, while the previous studies surveyed children or adolescents. We think this is a significant contribution since young adolescents still live with their parents, while most young adults live on their own and socialize with co-workers or classmates at university. Our study reveals that changes in gender norms induced by paternity leave reforms are persistent all the way through adulthood, even after leaving the family nest. Third, while accumulating evidence reveals that paternity leave reforms have had no strong effect on the labor supply decisions of the generation directly affected by the reforms (Andresen & Nix, forthcoming; Ekberg, Eriksson, & Friebel, 2013; Kleven, Landais, Posch, Steinhauer, & Zweimüller, forthcoming), our study reveals spillover effects on the next generation. Indeed, we demonstrate that the young men exposed to a father on paternity leave adopt counter-stereotypical behaviors in the labor market.

I. Paternity Leave Reforms in Six Countries

We evaluate the intergenerational spillover effects of paternity leave using reforms in six countries that were early adopters: Belgium (Jul. 2002), Denmark (Jan. 1995), France (Jan. 2002), Iceland (Jan. 2001), Norway (Apr. 1993) and Sweden (Jan. 1995). The Scandinavian countries introduced so-called "daddy months," earmarking one month of parental leave for fathers only. As a result, take-up increased to 60 percent in Denmark (Haataja, 2009), 80 percent in Iceland (Olafsson & Steingrimsdottir, 2020), 60 percent in Norway (Cools et al., 2015; Rege & Solli, 2013), and 80 percent in Sweden (Ekberg et al., 2013). Meanwhile, Belgium and France introduced a two-week long paternity leave exclusively for fathers, with a similarly high take-up rate from the start, up to 50 percent of fathers in Belgium (Fontenay & Tojerow, 2024), and 70 percent in France (Gosselin & Lepine, 2018).

Because the paternity leave introduction in our six countries dates back to the 1990's and early 2000's, the children born around the reform date have since entered adulthood. Indeed, our data shows that when taking the Implicit Association Test, our sample of respondents was on average 23 years old (Appendix Table A1). Our sample is therefore composed of young adults who, depending on their birthdate, were exposed or not to counter-stereotypical gender roles in their household during childhood and adolescence.

II. Measuring Gender Norms using the Implicit Association Test

Previous research has mostly relied on survey questions to measure gender role attitudes, asking for instance, whether women with young children should work or stay at home (Doepke & Kindermann, 2019). While survey questions might be able to measure long term changes across cohorts, they could fail to capture differences between individuals born in close proximity because of "social desirability biases," which push respondents to answer questions in a manner that make them look good to their peers, concealing their true opinions (Carlana et al., 2022).

The Implicit Association Test (IAT) is designed to neutralize social desirability bias by relying on subconscious mental associations. The IAT dates back to the seminal work of Greenwald, McGhee, and Schwartz (1998) and relies on test-takers' speed of response to capture the strength of their subconscious mental associations. Egloff and Schmukle (2002) show that IAT scores are, in fact, difficult to manipulate, and a growing number of studies reveal how the score strongly correlates with real-world psychological responses and economic decision-making (Bertrand, Chugh, & Mullainathan, 2005; Bursztyn, Chaney, Hassan, & Rao, 2024; Carlana, 2019; Glover, Pallais, & Pariente, 2017).

In our case, the participants who take part in the "Gender-Career" IAT on the Project Implicit website are asked to quickly sort words into categories: associating (i) "male" and "female" names with (ii) words related to "career" or "family." Because words and pairings are changed several times, the speed at which participants perform the task conveys information about implicit association. Our main outcome is a standardized index with higher values representing a stronger association between men and career and between women and family. Appendix Figure A1 reveals that younger cohorts who take the "Gender-Career" IAT display less stereotypical associations across all the countries considered.

The dataset that we use is publicly available from the Project Implicit website. We keep all respondents who live in the six countries that adopted paternity leave in the 1990's or early 2000's and who answered the online survey between 2016 and 2023. In addition to the test scores, IAT respondents are asked to provide information about gender, race, and nationality.[3] This provides us with important information to verify that respondents born right before and right after the reforms are similar in observables.

Most importantly, the rich dataset allows us to evaluate the effect of paternity leave across generations for men and women separately. We hypothesize that same-gender role models may have a stronger effect, so that young men might be more influenced by their fathers' counter-stereotypical behaviors. Our hypothesis relies on previous research showing strong influence of fathers on their sons' attitudes and behaviors, including smoking (Loureiro, Sanz-de-Galdeano, & Vuri, 2010), health and nutrition (Thomas, 1994), occupational choices (Lo Bello & Morchio, 2022), educational choices (Azam & Bhatt, 2015), and the decision to have children at a young age (Sipsma, Biello, Cole-Lewis, & Kershaw, 2010).

[3] The survey is provided in Dutch, English and French.

III. Empirical strategy

To measure the causal effects of paternity leave reforms across generations, we exploit the natural experiments provided by the policy changes in six countries. Using a Regression Discontinuity Design (RDD), we compare young adults born before and after the reform cutoffs:

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

where y is the IAT score of respondent i, born in month t around the reform cutoff c. f_r and f_l are unknown functions with parameter vectors γ_r and γ_l, capturing trends in the outcome of interest. We can interpret β as the estimated discontinuity for a respondent born just before versus just after the paternity leave reform. We do not observe paternity leave take-up in our data, so that our estimates should be interpreted as intent-to-treat (ITT) effects. We do know from previous studies in the six countries that the reforms led to high take-up, so that our ITT estimates should be driven by a meaningful share of the population.

Our estimates are produced using local polynomial regressions (Calonico, Cattaneo, & Titiunik, 2014a) and a uniform kernel (i.e. no weighting). We report heteroskedasticity-robust standard errors clustered at the month of birth (our running variable), following seminal work by Lee and Card (2008). We also report standard errors and "bias-aware" confidence intervals following the procedure suggested by Kolesár and Rothe (2018) for the RDD when the running variable only takes a moderate number of distinct values.[4]

Assuming no sorting in births around the reform date, we can interpret the estimated discontinuity at the cutoff as the causal effect of the paternity leave on the next generation's

[4] The Stata package "rdhonest" only allows for RDD estimation with a linear polynomial, so we provide results for this specific case along other robustness checks in Appendix Table A2.

gender role attitudes. This assumption is supported by the many papers that have previously studied the effects of the paternity leave reforms on parents in the six countries considered (see e.g. Avdic & Karimi, 2018; Cools et al., 2015; Ekberg et al., 2013; Olafsson & Steingrimsdottir, 2020; Rege & Solli, 2013).

We formally check for the absence of bunching in the number of observations around the reform cutoff in our dataset using a McCrary density test (2008). Our results in Table 1 suggest that parents did not manipulate the date of childbirth (e.g., through c-section or induced birth) to become eligible for paternity leave, and/or that exposure to paternity leave is unrelated to the likelihood of taking the gender-career IAT test. This is also confirmed by Appendix Figure A2, showing the density of the assignment variable (i.e. month of birth) and estimating the discontinuity at the reform cutoff using local-polynomial density estimators (Cattaneo, Jansson, & Ma, 2020). Furthermore, Table 1 reveals that predetermined characteristics of the IAT respondents are well balanced across the reform cutoff, whether it is the proportion of women, nonwhite individuals, Christians, or individuals with migration background. As suggested by previous evidence from the literature, the parents of respondents born before/after the reform were also equally likely to be working (see e.g. Andresen & Nix, forthcoming; Ekberg et al., 2013; Kleven et al., forthcoming).

Regarding the composition of our sample, respondents took the IAT test online voluntarily on the Project Implicit website. The last panel of Table 1 reveals that most respondents took the IAT because of university or work assignment (62 percent). Others found out about the IAT on the news (18 percent) or through friends (20 percent) and decided to take the test as a result. Most importantly, the reasons to take the test are perfectly balanced around the reform cutoff, so that we can rule out that IAT participants born after the reform were keener to participate in a survey on gender norms.

IV. Effects of Paternity Leave on the Next Generation

A. Gender Role Attitudes

We now report the results on the effect of paternity leave on the gender role attitudes of the next generation. We start by pooling together IATs for all the respondents in the six countries. Figure 1 shows the discontinuity in test scores at the reform cutoff (vertical bar) for men in Panel A and women in Panel B. We use a bandwidth that spans a couple of years around the reform to be able to check visually for the absence of season of birth effects that could influence the result. We clearly observe that men born right after the reform, and potentially exposed to a father on paternity leave, display less gender-stereotypical attitudes. In other words, male respondents affected by the paternity leave reform are less likely to associate women with home-making and men with career. Meanwhile, we do not detect any change in attitudes for female respondents.

In Table 2, we estimate the discontinuity at the reform cutoff using a polynomial of order 0 and a sample that includes 24 months before and after the reform threshold. We choose this one as our main specification because our graphical exploration revealed no sign of a trend in the outcome variable in the small window around the reform. We find that young men born right after the paternity leave reform are less likely to associate men with career and women with family. Confirming our visual inspection, we find no effect on young women. The magnitude of the effect for men is a reduction of 0.21 standard deviations in stereotypical association. Appendix Figure A1 is useful to provide a sense of the profound changes that the paternity leave reforms produced in the next generation. The estimated difference between adults born only one month apart, but exposed or not to a father on paternity leave, is similar to five decades of reduction in gender-stereotypical attitudes measured by the IAT in countries like Canada (-0.18) or the United States (-0.2). It is also of the same order of magnitude than the differences measured between cohorts born in 1960s in gender-equality pioneer Sweden and France (-0.17) or Belgium (-0.24).

Table 2 also provides effects for each country separately. Two countries stand out with particularly large effects on men by as much as 0.496 standard deviations in Belgium and 0.325 standard deviations in Sweden, both effects being highly statistically significant. We also notice a more modest negative effect of 0.181 standard deviations in Denmark, statistically significant at the 10 percent level. In the remaining countries, the effect on men is also negative and similar in magnitude to the overall effect (except in Norway), but not statistically significant at conventional levels. We also find consistently no effect on women in each of the six countries considered.

As mentioned in section III, most respondents participate in the IAT because of a university or work assignment. As such, we can expect that the type of workers who have been asked by their employer to take the test might differ from the general population. In fact, Appendix Table A1 reveals that 70 percent of the participants attended tertiary education, while only 6 percent have low educational attainment (primary or lower secondary school). Our sample of respondents therefore sensibly differs from the typical educational attainment of the population in the six countries considered. We reweight our sample so as to better reflect the educational attainment of the general population, the goal being to make our sample closer to being representative of the overall population.[5] Results reported in Panel B of Table 2 suggest that the effect on male IAT respondents is, if anything, larger when reweighing the sample. The effects remain essentially unchanged in Belgium and Sweden, and are now larger and statistically significant in France (at the 5 percent level).

[5] According to OECD data for the six countries considered in our study, 17 percent of the general population has primary or lower secondary education, 39 percent are high school graduates, and 44 percent reached tertiary education.

Another concern related to the availability of the IAT online is that some people decide to take the test on their own and, therefore, self-select into the project. As mentioned in section III, we do not detect bunching in the number of participants born after the reforms. In addition, pre-determined characteristics are perfectly balanced across the reform threshold. However, one might be concerned with unobserved characteristics changing the composition of respondents on either side of the reforms cutoff. For instance, individuals more concerned with gender equality might hear about the test in the news or through friends and decide to take the test. To alleviate this concern, we remove from the sample all the participants who declared that they took the test upon a friend's recommendation or after reading about it in the news. The "not self-selected sample" in Appendix Table A2 reveals that the effects are highly similar to our main specification using the complete sample of IAT respondents. To further dissipate concerns about self-selection, we investigate other IATs related to racial prejudice, sexual orientation or physical characteristics, also publicly available on the Project Implicit website. The intuition is that the paternity leave reforms, by providing a counter-stereotypical example of a father as homemaker, should not impact the score of respondents in other IATs that are not related to gender norms. If they did, then one could suspect that other confounders might be driving the effects. Appendix Figure A3 reveals a precisely estimated 0 difference on the race IAT taken by young adults born right after and right before the paternity leave reforms. This is confirmed in other IATs related to sexual orientation, weight, and age in Appendix Table A3.

B. Occupational Choice

In the previous subsection, we demonstrate that paternity leave reforms had a long-lasting impact on the next generation's gender role attitudes, most crucially among men. We now turn to examining whether this change in attitudes affected real-life decisions. In particular, we investigate whether the young men exposed to a father on paternity leave adopted less gender-stereotypical behaviors when entering the labor market. We argue after Charles and Grusky

(2005) that studying policies that could reduce "occupational ghettos" is important since gender occupational segregation has been found to be very persistent over time, even in countries where other types of gender inequality have lessened (e.g. labor force participation or wage gap). Most crucially, while women have increasingly entered male-dominated occupations over the past decades (Busch, 2020; Mandel, 2012), the presence of men in typically female-dominated occupations is still very scarce (Torre, 2014). Conservative attitudes and values seem to be a crucial barrier (Davis & Greenstein, 2009; Irmert, 2024), and therefore, the change in gender norms brought about by paternity leave reforms may be a new driving force pushing men into counter-stereotypical roles. In fact, previous research in sociology suggests that young men with more egalitarian attitudes are more likely to aspire toward more female-dominated occupations (Baird, 2012; Correll, 2001).

We use data from the survey on Statistics on Income and Living Conditions (SILC), which is probably the best effort within the European Union to collect data related to education, labor market and income. Most importantly for our study, the questionnaire is harmonized across countries and highly similar over time. We aggregate cross-sectional survey waves from 2006 to 2020 and keep those respondents born within 5 years of the paternity leave reform cutoff (e.g. for Denmark, where the reform took place in Jan. 1995, we keep respondents born between 1990 and 1999). Since we are primarily interested in occupational segregation, we further restrict the sample to adult respondents (i.e. aged at least 18 years old). Because of this last restriction, our sample only includes three of the six countries, those that adopted paternity leave in the 1990s: Denmark, Norway and Sweden.[6]

[6] The last survey wave of SILC that contains the quarter of birth is 2020 for Belgium and France, and 2018 for Iceland. Because the paternity leave reforms took place in the early 2000s in those countries, respondents are not yet 18 when taking the survey.

Appendix Table A4 provides descriptive statistics on the sample of 45,950 SILC respondents born within a five-year bandwidth around the reform cutoff, aged over 18 at the time of the survey, and residing in Denmark, Norway or Sweden. The average respondent is 21.5 years old, and very few are married or have children. One out of five respondents has completed post-secondary education, while four out of five are working.[7] The SILC questionnaire includes a question on occupation following the ISCO-08 classification, with a total of 43 different occupations. We compute the share of women in each occupation. Appendix Table A5 reveals that women are disproportionately represented among personal care workers (87%), keyboard clerks (86%), cleaners/helpers (81%) and health professionals (80%). Our main outcome variable is a dummy that takes on the value 1 if a SILC respondent is working in one of the 17 occupations in which the majority of workers are women, and 0 otherwise.

Our empirical strategy follows again Equation (1), except that the running variable is now the quarter of birth (due to the SILC anonymization strategy). In a similar exercise than before, Figure 2 shows the probability to work in a female-dominated occupation in three-month bins, for cohorts born during the five years before and after the reform cutoff. Panel A of Figure 2 reveals that the share of men working in female-dominated occupation evolves around 45 percent in the 20 quarters before the paternity leave reforms. Then, we observe a sharp discontinuity at the reform cutoff, with an estimated increase of 4.5 percentage points in the fraction of men working in a female-dominated occupation (Panel A of Table 3). The effect is particularly large for Sweden (Panel B of Table 3), precisely the country in which we measured the largest change in gender role attitudes after the paternity leave reform in 1995. In accordance with our previous results on gender norms, we also find both in Figure 2 (Panel B) and the second column in Table 3 (Panel A) that there is no effect on women. We check (in Appendix

[7] The probability to work takes the value 1 if the respondent reports wage earnings higher than 0.

Table A6) that these results are not driven by labor force participation or human capital decisions, and we find that there is no change at the reform cutoff in the probability to work or having completed post-secondary studies.

In a final exercise, we further distinguish between low-skilled and high-skilled occupations by mapping the ISCO-08 major groups to skill levels using the correspondence table of the International Labour Office (2012). We split our sample between SILC respondents working in low-skilled occupations (levels 1 and 2) and high-skilled occupations (levels 3 and 4). The second row of Panel A in Table 3 reveals that treated men are more likely to work in a female-dominated occupation classified as highly skilled (e.g. health or teaching professionals). We find no effect on low-skilled occupations.

C. Robustness Checks

In this subsection, we offer several robustness checks related to the model specification, as well as the issue of seasonality, for both the IAT results on gender role attitudes and the SILC results on female-dominated occupations.

Starting with the sensitivity of our results to model specification, we provide evidence in Appendix Tables A2 (for the IAT sample) and A7 (for the SILC sample) that our results are robust to using different polynomial orders (linear or quadratic), as well as different bandwidth sizes. Most importantly, we show that our results are robust to using the data-driven procedure of Calonico, Cattaneo, and Titiunik (2014b) to choose the bandwidth (“BW = CCT”). The robustness of our results is also confirmed when computing “honest confidence intervals,” as suggested by Kolesár and Rothe (2018) when dealing with a discrete running variable (“Honest 95% CI”).

Next, we want to rule out that our results might suffer from seasonality issues, since the previous literature has highlighted how the season of birth might be related to later life outcomes because of school starting age or parental selection (Buckles & Hungerman, 2013). In our specific context, we do not anticipate the seasonal effects to interfere strongly with our estimates since the cutoff is different in two of the countries where we measure the largest effects, namely Belgium (July) and Sweden (January). However, we remain cautious and use two different strategies to test for the presence of season of birth confounding effects. First, we change our main specification to account for season of birth effects. We add calendar-month-of-birth fixed effects to equation (1) for the IAT sample, and calendar-quarter-of-birth fixed effects for the SILC sample. Panel A of Appendix Table A8 reveals that the effect on the IAT score is very similar to our baseline estimates, while Appendix Table A9 also confirms the robustness of our findings on occupations. Second, we augment our RDD specification with a Difference-in-Differences strategy to net out any potential seasonality in the outcome variables (similar to Dustmann and Schönberg (2012)). Our coefficient of interest is now the discontinuity at the reform cutoff within a tight six-month bandwidth, compared to non-reform years. In the Belgian case, for instance, we estimate the effect at the reform cutoff in July 2002, compared to the seasonal discontinuity in July during pre-reform years from 1997 to 2001. Panel B of Appendix Table A8 (Table A9) shows, once again, highly similar effects at the cutoff during the reform year ("Cutoff * Reform year") for the IAT sample (for the SILC sample).

Another way to rule out the season of birth effects, and to show that our RDD strategy is truly capturing exceptional circumstances at the reform cutoff, is to compare our estimates to placebo estimates during pre-reform years. To do so, we artificially move the cutoff to the same calendar month (quarter) in each of the 25 years before the actual paternity leave reforms, and show the distribution of the estimates in Appendix Figure A4 for the IAT sample (Appendix Figure A5 for the SILC sample). One can see in Panel A of Appendix Figures A4 and A5 that, for male respondents, the placebo estimates are centered around 0, while our estimate for the actual reform year (represented by the red dashed line) is completely off on its own and far away from the rest of the distribution. As such, we are truly capturing a special event during the reform year that does not overlap with typical deviations measured in non-reform years.

V. Mechanisms

In this last section, we explore the mechanisms at play and discuss how young adults born only a few months apart could display very different gender role attitudes. We hypothesize that the fathers eligible for paternity leave have acted as a role model for their sons, who, later in life, consider both parents as homemakers and display less stereotypical attitudes when taking the IAT. This assumes, of course, that the reforms have had an effect beyond the paternity leave period that takes place immediately after birth, and that the children of eligible fathers have had the chance to be exposed to a different role model. Appendix Table A10 surveys the previous literature and reveals that the reforms in the six countries considered have had a long lasting impact on the division of tasks in the households. For instance, Fontenay and Tojerow (2024) find that, up to 10 years after childbirth, post-reform fathers in Belgium dedicated twice as much time to their children, compared to the control ones. We reviewed the literature and found evidence of long term changes in the household division of labor in nine different countries where paternity leave was introduced.

We seek to confirm these previous findings and establish cross-country evidence that paternity leave reforms affect the households' division of labor in the long run. To do so, we use the European Working Conditions Survey (EWCS), which, despite not being conducted to collect data on household chores and childcare, has several advantages. First, it is a nationally representative survey conducted in five of the countries considered in this study: Belgium, Denmark, France, Norway and Sweden (only Iceland is missing). Second, the questionnaire is harmonized and asks questions about how often interviewees perform various activities, including childcare and housework. Finally, the 2005 EWCS collected data on fathers who had children before and after the reforms in the countries of our sample. Using a regression discontinuity design, we compare fathers who had a first child (and only child by the time they answered the survey) around the reforms' cutoff in each country. Unfortunately, we only have the birth year of the child, so that our estimates will rely on breaks in long run trends. Panel A of Appendix Figure A6 reveals that across the five countries considered, the fathers who were eligible for paternity leave are 14 percentage points more likely to report dedicating time to housework weekly. Panel B does not suggest any effect on childcare time, which might stem from the fact that most of the fathers of pre-reform children were already dedicating time each week to them, so that the margin of improvement that the survey can capture was rather thin.

We conclude from this new analysis that the impact of the paternity leave reforms goes far beyond the leave period. In particular, the evidence suggests that the fathers eligible for paternity leave devoted more time to household work persistently. We believe that, by offering a counter-stereotypical example to their sons during all their childhood, those fathers eligible for paternity leave have contributed to the change in gender norms that we observe in the next generation.

VI. Conclusions

We study the effect of paternity leave policies on attitudes about gender roles in the next generation. We follow a Regression Discontinuity Design where the running variable is the month of birth, and study the results in an Implicit Association Test about gender, for young adults born shortly before and after the introduction of paternity leave in six different European countries. We find that men exposed to fathers who were eligible for paternity leave display significantly less gender-stereotypical attitudes as adults. This effect is similar in five out of the six countries in our sample (about 0.2 of a standard deviation), while we find no effect for women. We also provide suggestive evidence that paternity leave reforms persistently increased fathers' involvement in household chores, and conjecture that the main mechanism behind our results is a role-model effect, such that boys who grow up with a father who is more involved in the home throughout their childhood develop less gender-stereotypical attitudes.

Building on the observed change in gender role attitudes, we use survey data to examine how this translates into counter-stereotypical behaviors in the labor market. We find that male respondents born just after the paternity leave reforms in Denmark, Norway, and Sweden are 4.5 percentage points more likely to work in a female-dominated occupation, particularly in Sweden (and in high-skilled jobs), underscoring the potential of paternity leave reforms to mitigate gender segregation in the labor market.

Our results highlight the potentially far-reaching effects of policies that affect gender roles in households with children, if attitudes are shaped by parental behaviors during childhood. It remains to be seen the extent to which the changes in gender role attitudes, as well as the counter-stereotypical behaviors that we observe in the labor market, will translate into smaller child penalties (and thus gender gaps) in earnings in the next generation.

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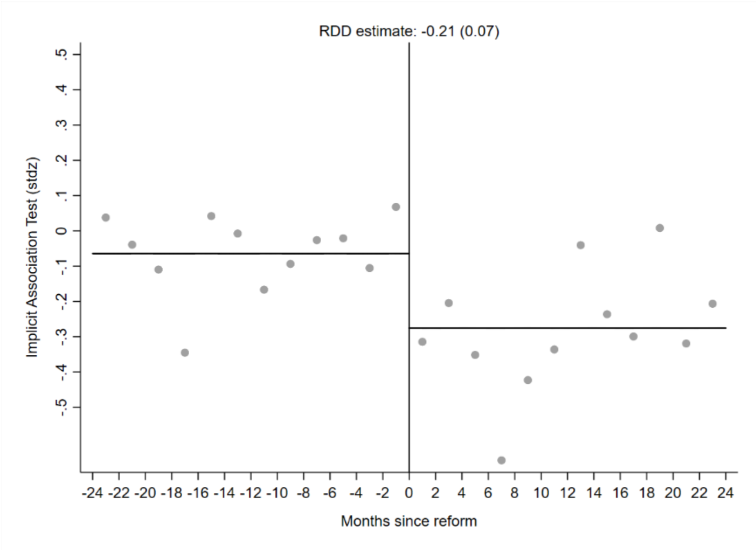
Panel A: Sample of Male Respondents



Panel B: Sample of Female Respondents

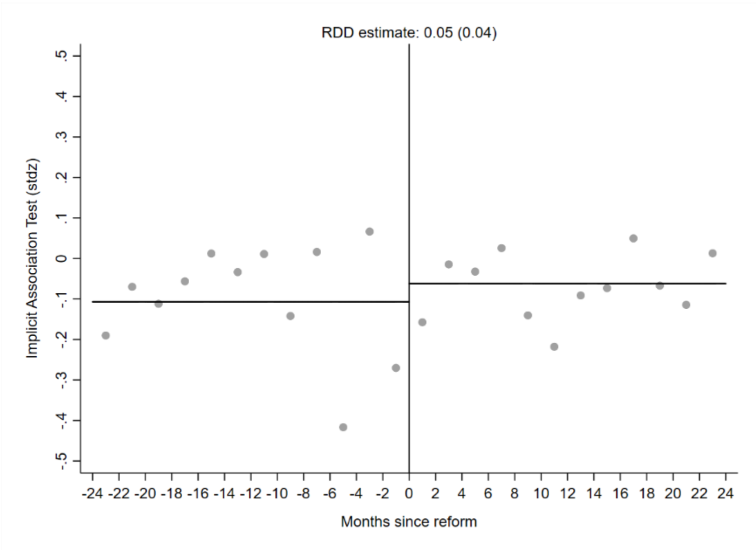
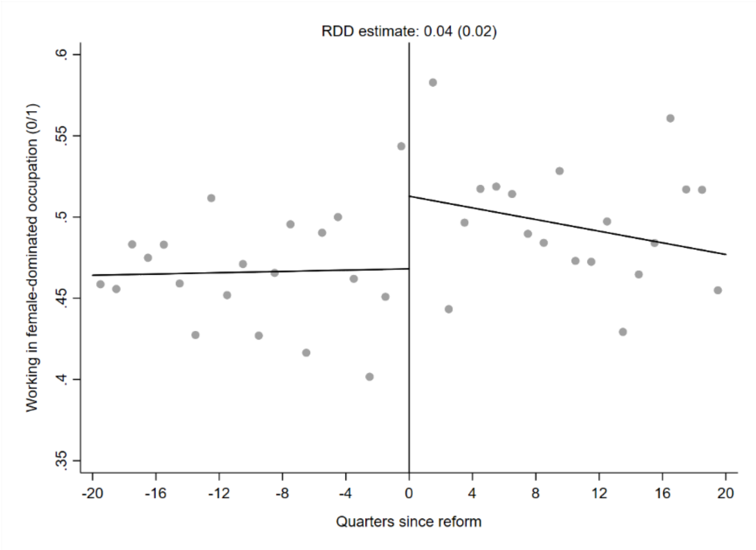


Figure 1: Change in Implicit Association Test Score at the Reform Cutoff.

Notes: Average IAT score in four-month bins. The vertical bar corresponds to the reform cutoff, normalized to 0 in each of the six countries. The horizontal bars on each side of the cutoff are from local polynomial regression of order 0. The RDD estimate reported on the top of the graph corresponds to coefficient β in equation (1). Sample of respondents to the "Gender-Career" IAT on the Project Implicit website who live in Belgium, Denmark, France, Iceland, Norway and Sweden.

Panel A: Sample of Male Respondents



Panel B: Sample of Female Respondents

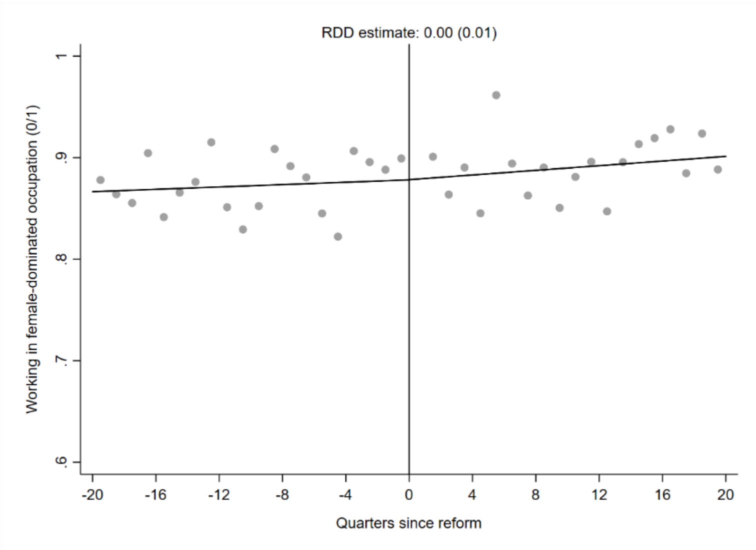


Figure 2: Change in the Probability to Work in a Female-Dominated Occupation at the Reform Cutoff.

Notes: Share of SILC respondents aged over 18, who work in a female-dominated occupation (ISCO-08 classification) in three-month bins. The vertical bar corresponds to the reform cutoff, normalized to 0 in each of the three countries (Denmark, Norway and Sweden). The trends on each side of the cutoff are from local polynomial regression of order 1. The RDD estimate reported on the top of the graph corresponds to coefficient β in equation (1).

Table 1: Balance in Covariates

	Coef. (SE)	Mean	Nb. observations
McCrary density test			
Discontinuity at reform cutoff (log diff.)	-0.010		
	(0.058)		
Pre-determined characteristics			
Female (0/1)	0.018	0.605	3342
	(0.016)		
Nonwhite (0/1)	0.002	0.131	3202
	(0.013)		
Christian (0/1)	0.009	0.339	3315
	(0.017)		
Migration background (0/1)	-0.021	0.246	3340
	(0.014)		
Family structure during youth			
Primary caregiver = Mother (0/1)	-0.013	0.764	3348
	(0.018)		
Primary caregiver = Mother + Working (0/1)	-0.003	0.569	3348
	(0.018)		
Secondary caregiver = Father (0/1)	-0.013	0.665	3348
	(0.018)		
Secondary caregiver = Father + Working (0/1)	-0.016	0.650	3348
	(0.018)		
Reasons to take test			
Assignment from school/work (0/1)	0.021	0.617	1624
	(0.024)		
Recommendation of friend or co-worker (0/1)	-0.015	0.204	1624
	(0.019)		
News, internet, other (0/1)	-0.006	0.179	1624
	(0.015)		

Notes: The first panel titled "McCrary density test" checks for the absence of bunching in the number of births after the reform cutoff using a McCrary density test (2008). The rest of the table reports in the first column RDD estimates from local polynomial regression of order 0 and corresponding to coefficient β in equation (1). Standard errors (reported in parentheses) are clustered at the birth of month level (i.e. running variable). The second and third column report the mean and number of observations, respectively. Sample of respondents to the "Gender-Career" IAT on the Project Implicit website who live in Belgium, Denmark, France, Iceland, Norway and Sweden.

Table 2: Effect of Paternity Leave Reforms on Gender Norms

	Standardized Implicit Association Test					
	Panel A: Unweighted sample			Panel B: Weighted sample		
	Both	Men	Women	Both	Men	Women
	(1)	(2)	(3)	(4)	(5)	(6)
All countries						
RDD coef.	-0.042	-0.211 ***	0.045	-0.061	-0.286 ***	0.058
(SE)	(0.033)	(0.065)	(0.044)	(0.040)	(0.074)	(0.048)
Nb. observations	3348	1159	2183	3282	1135	2142
Belgium						
RDD coef.	-0.102	-0.496 ***	0.090	-0.061	-0.431 **	0.099
(SE)	(0.097)	(0.169)	(0.108)	(0.106)	(0.173)	(0.117)
Nb. observations	445	143	301	433	139	294
Denmark						
RDD coef.	-0.038	-0.181 *	0.053	-0.009	-0.129	0.077
(SE)	(0.057)	(0.106)	(0.079)	(0.073)	(0.119)	(0.100)
Nb. observations	774	299	475	764	296	468
France						
RDD coef.	-0.079	-0.161	-0.038	-0.156 **	-0.289 **	-0.074
(SE)	(0.056)	(0.123)	(0.066)	(0.069)	(0.137)	(0.085)
Nb. observations	1035	335	695	1019	330	684
Iceland						
RDD coef.	0.038	-0.351	0.219	0.066	-0.480	0.269
(SE)	(0.196)	(0.316)	(0.230)	(0.212)	(0.333)	(0.260)
Nb. observations	82	28	54	77	24	53
Norway						
RDD coef.	0.055	0.144	-0.006	-0.014	0.064	-0.074
(SE)	(0.110)	(0.203)	(0.130)	(0.128)	(0.231)	(0.142)
Nb. observations	320	111	209	312	109	203
Sweden						
RDD coef.	-0.048	-0.325 ***	0.104	-0.019	-0.395 **	0.177
(SE)	(0.078)	(0.119)	(0.107)	(0.091)	(0.151)	(0.113)
Nb. observations	692	243	449	677	237	440

Notes: The table reports RDD estimates from local polynomial regression of order 0 and corresponding to coefficient β in equation (1). Standard errors (reported in parentheses) are clustered at the birth of month level (i.e. running variable). Bandwidth of 24 months around reform cutoff. The left panel reports estimates for an unweighted sample, while the right panel is for a reweighted sample reflecting educational attainment of the general population. Sample of respondents to the "Gender-Career" IAT on the Project Implicit website who live in Belgium, Denmark, France, Iceland, Norway and Sweden.

Table 3: Effect of Paternity Leave Reforms on Occupational Choice

	Female-dominated occupation (0/1)	
	Men	**Women**
	RDD coef. (SE) Obs.	RDD coef. (SE) Obs.
Panel A - All countries		
All occupations	0.045 **	0.000
	(0.018)	(0.013)
	13573	11643
High-skilled occupations	0.106 ***	0.044
	(0.039)	(0.031)
	3082	3290
Low-skilled occupations	0.026	-0.014
	(0.020)	(0.014)
	10481	8353
Panel B - Country breakdown		
Denmark	0.011	-0.038
	(0.041)	(0.032)
	2990	2395
Norway	0.035	0.005
	(0.025)	(0.018)
	6425	5560
Sweden	0.087 **	0.018
	(0.035)	(0.024)
	4158	3688

Notes: The table reports RDD estimates from local polynomial regression of order 1 and corresponding to coefficient β in equation (1). Robust standard errors are reported in parentheses. Bandwidth of 20 quarters around the reform cutoff. The running variable is the quarter of birth. Sample of respondents to the SILC survey, aged 18 or above from Denmark, Norway and Sweden.

Appendix

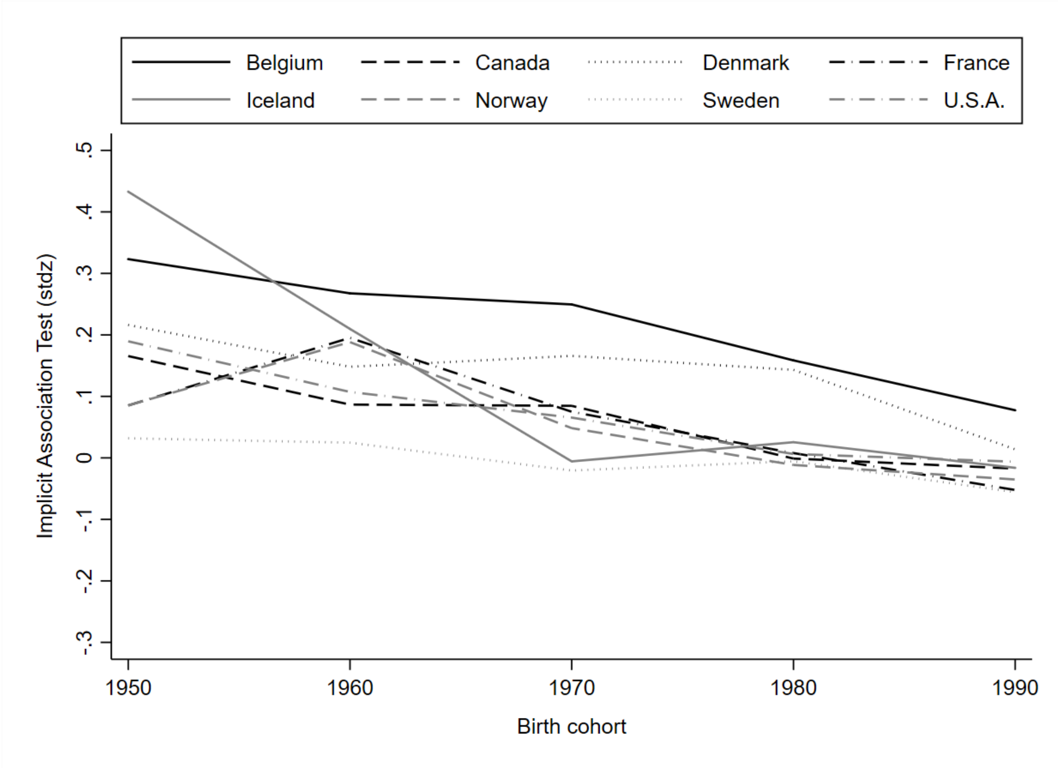


Figure A1: Implicit Association Test Score across Birth Cohorts

Notes: Lower score on the standardized IAT indicates a lower stereotypical association between women and family and between men and career. Sample of respondents to the "Gender-Career" IAT on the Project Implicit website who live in Belgium, Canada, Denmark, France, Iceland, Norway, Sweden and the United States.

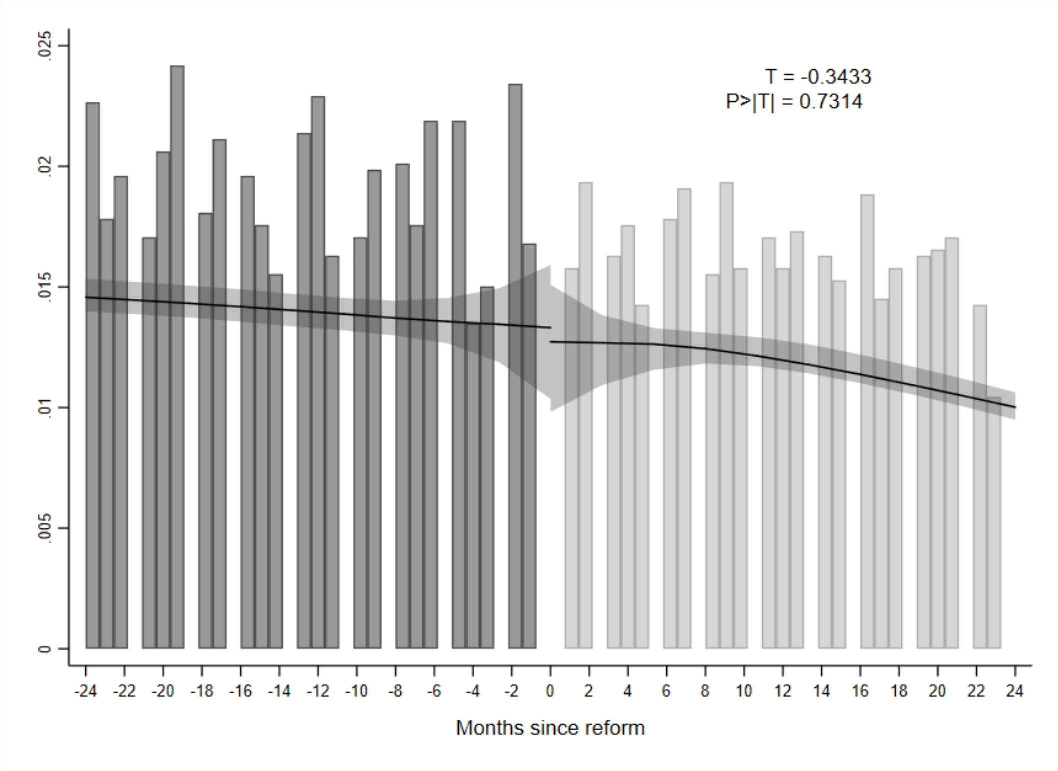
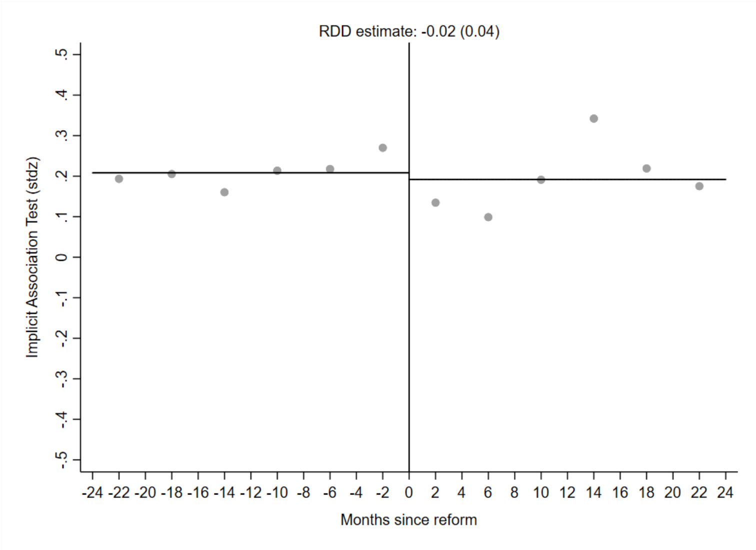


Figure A2: Manipulation Testing using Local-polynomial Density Estimation

Notes: The figure displays the density of the assignment variable (i.e. month of birth) around the cutoff, estimated using local-polynomials, as proposed by Cattaneo et al. (2020). The authors also suggest a novel manipulation test, which is reported on the graph with the corresponding p-value.

Panel A: Sample of Male Respondents



Panel B: Sample of Female Respondents

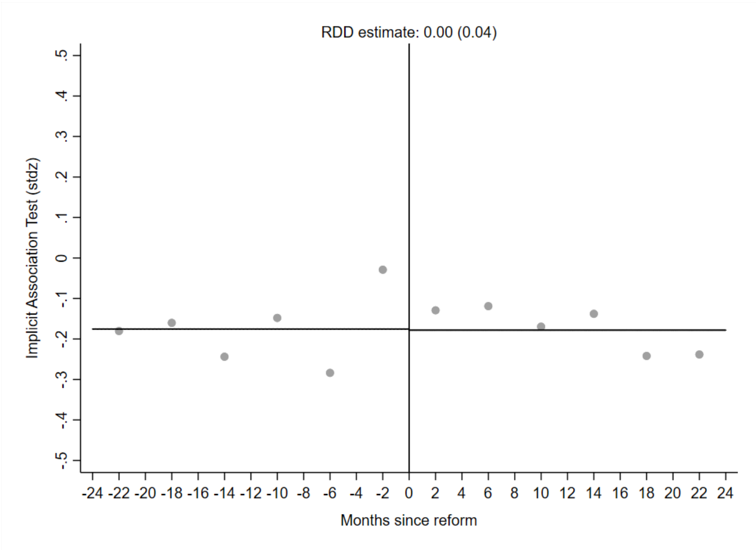
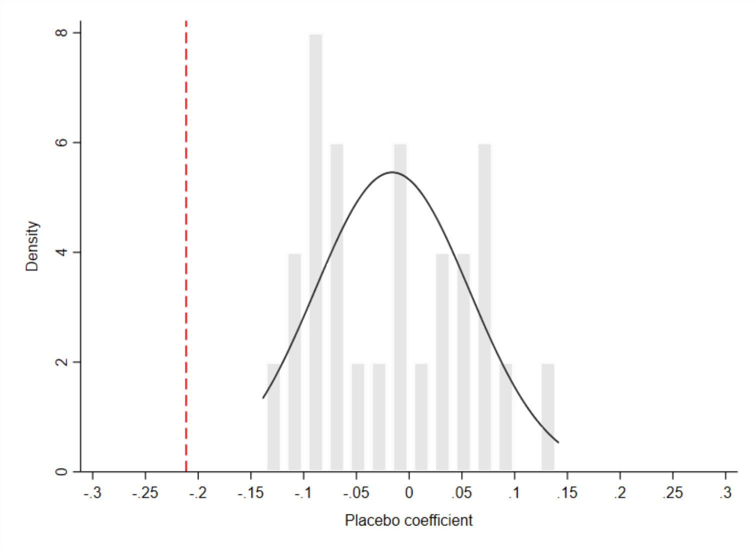


Figure A3: Change in Race IAT at the Reform Cutoff.

Notes: Average IAT score in four-month bins. The vertical bar corresponds to the reform cutoff, normalized to 0 in each of the six countries. The horizontal bars on each side of the cutoff are from local polynomial regression of order 0. The RDD estimate reported on the top of the graph corresponds to coefficient β in equation (1). Data collected via an IAT measuring the subconscious association between race (black - white), and pleasant or unpleasant words. Sample of respondents who live in Belgium, Denmark, France, Iceland, Norway and Sweden.

Panel A: Sample of Male Respondents



Panel B: Sample of Female Respondents

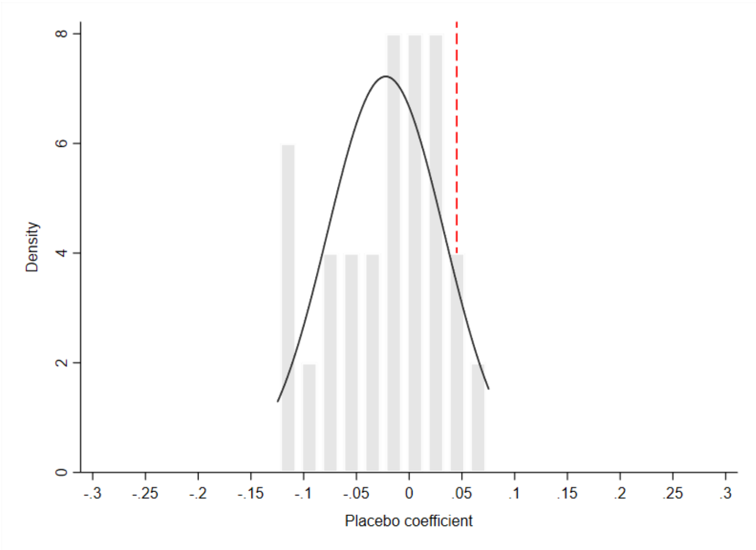
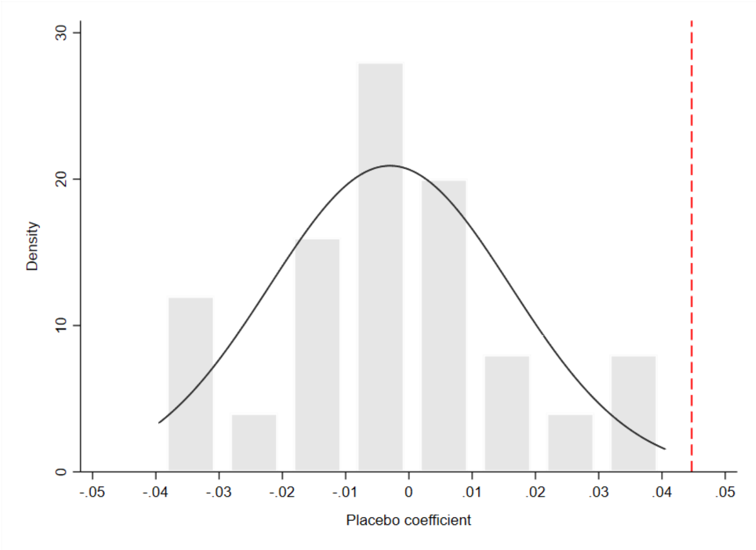


Figure A4: Change in IAT Score at Placebo Cutoffs.

Notes: Distribution of RDD estimates at placebo cutoffs during the 25 years before the paternity leave reform in each country (e.g. for Norway, the cutoff is moved to April in each year between 1967 to 1992, before the actual reform in 1993). The red dashed line corresponds to the RDD estimate at the true reform cutoff. Sample of respondents who live in Belgium, Denmark, France, Iceland, Norway and Sweden.

Panel A: Sample of Male Respondents



Panel B: Sample of Female Respondents

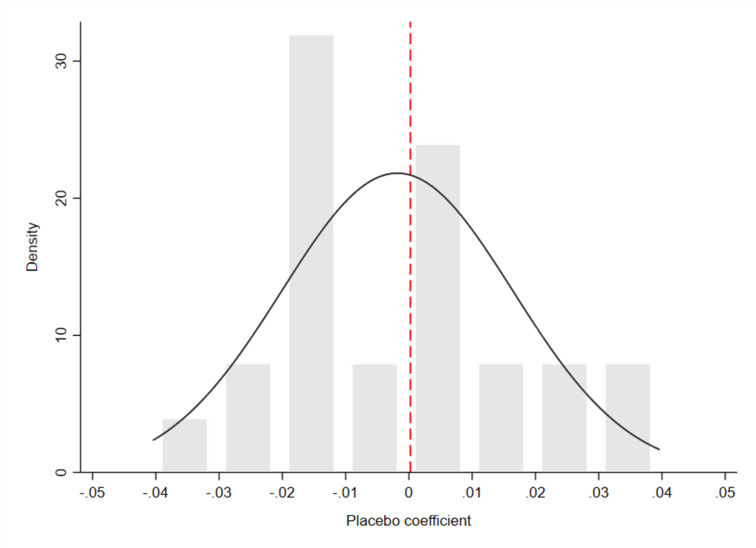
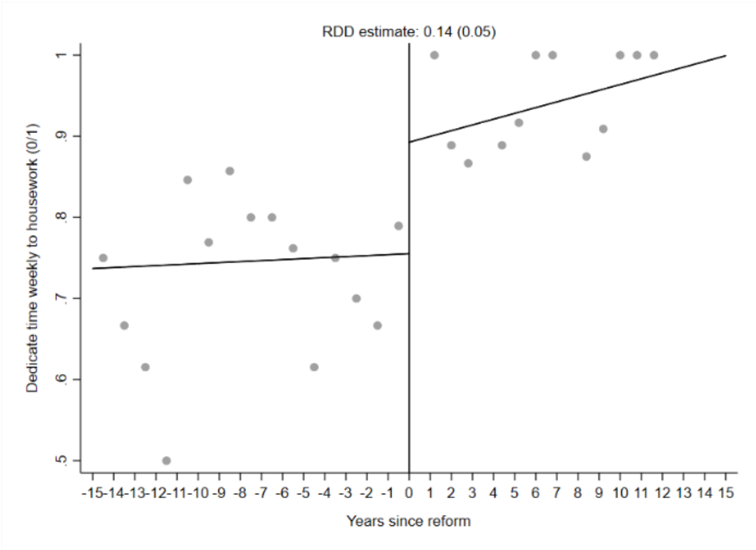


Figure A5: Change in Probability to Work in a Female-dominated Occupation at Placebo Cutoffs.

Notes: Distribution of RDD estimates at placebo cutoffs during the 25 years before the paternity leave reform in each country (e.g. for Norway, the cutoff is moved to April in each year between 1967 to 1992, before the actual reform in 1993). The red dashed line corresponds to the RDD estimate at the true reform cutoff. Sample of respondents to the SILC survey from Denmark, Norway and Sweden, aged at least 18 years old.

Panel A: Housework



Panel B: Childcare

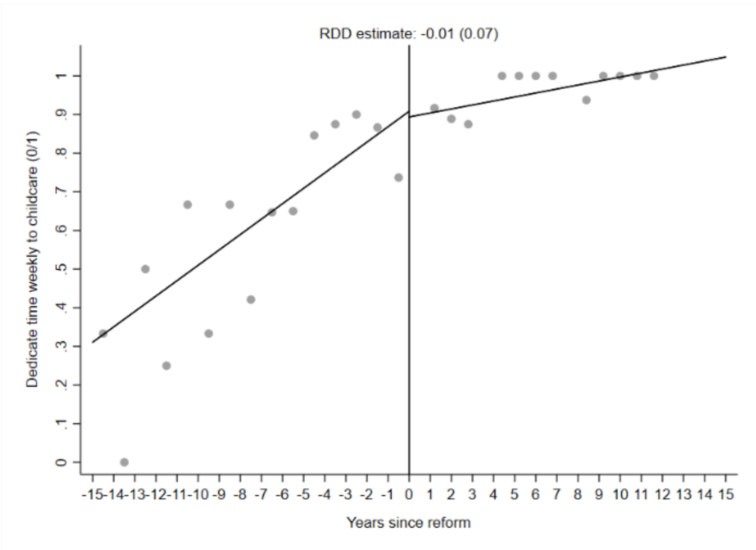


Figure A6: Effect of Paternity Leave Reforms on Fathers' Housework and Childcare.

Notes: Share of fathers dedicating time weekly to each activity in one-year bins. The vertical bar corresponds to the reform cutoff, normalized to 0 in each of the five countries. The horizontal bars on each side of the cutoff are from local polynomial regression of order 1. The running variable if the birth year of their first child. Data are from the 2005 European Working Conditions Survey. The exact question is "How often are you involved with cooking and housework? Caring for and educating your children? The outcome takes on a value of 1 if the answer is "Everyday for 1 hour or more," "Everyday or every second day for less than one hour," "Once or twice a week" and 0 otherwise. Sample of male respondents who live in Belgium, Denmark, France, Norway and Sweden, and had a first child around the reform cutoff and no additional children by the time they answered the survey.

Table A1: Descriptive Statistics on the Sample of IAT Respondents

	IAT sample 6 countries	OECD data 6 countries	Belgium	Denmark	France	Iceland	Norway	Sweden
Age	22.92		19.22	25.82	20.05	19.73	26.34	25.17
	(3.50)		(1.43)	(2.09)	(1.59)	(1.69)	(2.20)	(2.15)
Primary / lower secondary edu. (0/1)	0.06	0.17	0.22	0.01	0.07	0.29	0.01	0.01
	(0.24)		(0.42)	(0.07)	(0.26)	(0.45)	(0.10)	(0.11)
High school graduate (0/1)	0.24	0.39	0.60	0.11	0.27	0.34	0.06	0.17
	(0.42)		(0.49)	(0.31)	(0.44)	(0.48)	(0.25)	(0.38)
Tertiary education (0/1)	0.70	0.44	0.18	0.89	0.66	0.38	0.93	0.82
	(0.46)		(0.38)	(0.32)	(0.47)	(0.49)	(0.26)	(0.39)
Children (0/1)	0.03		0.02	0.04	0.02	0.01	0.05	0.02
	(0.16)		(0.14)	(0.20)	(0.13)	(0.11)	(0.21)	(0.15)

Notes: The table reports the mean and standard deviation (in parentheses) for different outcomes. IAT Sample correspond to the sample of respondents to the "Gender-Career" IAT on the Project Implicit website who live in Belgium, Denmark, France, Iceland, Norway and Sweden. OECD data are population averages from the OECD.

Table A2: Robustness Checks – IAT Sample

	Standardized Implicit Association Test		
	Both	Men	Women
	(1)	(2)	(3)
Main specification (Poly. = 0, BW=24)			
RDD coef.	-0.042	-0.211 ***	0.045
(SE)	(0.033)	(0.065)	(0.044)
Nb. observations	3348	1159	2183
Not self-selected sample			
RDD coef.	-0.045	-0.216 ***	0.042
(SE)	(0.037)	(0.070)	(0.048)
Nb. observations	2946	1015	1927
Poly. = 1			
RDD coef.	-0.049	-0.344 ***	0.091
(SE)	(0.065)	(0.121)	(0.099)
Nb. observations	3348	1159	2183
Poly. = 2			
RDD coef.	0.064	-0.473 **	0.319 **
(SE)	(0.094)	(0.185)	(0.128)
Nb. observations	3348	1159	2183
BW = CCT			
RDD coef.	-0.045	-0.270 ***	0.041
(SE)	(0.039)	(0.085)	(0.054)
Nb. observations	2519	640	1656
BW = 12			
RDD coef.	-0.060	-0.302 ***	0.061
(SE)	(0.046)	(0.088)	(0.073)
Nb. observations	1689	582	1105
BW = 18			
RDD coef.	-0.045	-0.220 ***	0.041
(SE)	(0.040)	(0.080)	(0.055)
Nb. observations	2519	861	1656
BW = Honest; Poly. = 1			
RDD coef.	-0.067	-0.306 ***	0.091
(SE)	(0.061)	(0.095)	(0.071)
Honest CI	[-0.193 ; 0.059]	[-0.503 ; -0.109]	[-0.056 ; 0.238]
Nb. observations	4645	2011	3174

Notes: The table reports RDD estimates based on equation (1) and from local polynomial regression with different polynomial orders. Standard errors (reported in parentheses) are clustered at the birth of month level (i.e. running variable). The lower panels also report estimates with different bandwidths around the reform cutoff, including the data-driven bandwidth suggested by the procedure of Calonico et al. (2014b) and the "honest confidence intervals" from Kolesár and Rothe (2018). Sample of respondents to the "Gender-Career" IAT on the Project Implicit website who live in Belgium, Denmark, France, Iceland, Norway and Sweden. The "not self-selected sample" excludes respondents who decided to take the test after hearing about it on the news or following a friend's recommendation.

Table A3 Effect of Reforms on Other Implicit Association Tests

	Standardized Implicit Association Test		
	All sample	Men	Women
Race (black - white)			
RDD coef.	-0.009	-0.016	-0.003
(SE)	(0.028)	(0.038)	(0.038)
Nb. observations	4670	2205	2453
Sexuality (gay - straight)			
RDD coef.	0.033	0.050	0.065
(SE)	(0.040)	(0.073)	(0.046)
Nb. observations	2761	1000	1755
Age (young - old)			
RDD coef.	-0.068	0.010	-0.086
(SE)	(0.053)	(0.089)	(0.061)
Nb. observations	1656	578	1075
Weight (fat - thin)			
RDD coef.	-0.051	-0.059	-0.050
(SE)	(0.040)	(0.075)	(0.058)
Nb. observations	2015	671	1343

Notes: The table reports RDD estimates from local polynomial regression of order 0 and corresponding to coefficient β in equation (1). Standard errors (reported in parentheses) are clustered at the birth of month level (i.e. running variable). Bandwidth of 24 months around reform cutoff. Data collected via four IATs measuring the subconscious association between age (young - old), race (black - white), sexuality (gay – straight), weight (fat - thin) and pleasant or unpleasant words. Sample of respondents who live in Belgium, Denmark, France, Iceland, Norway and Sweden.

Table A4: Descriptive Statistics on the Sample of SILC Respondents

	All	Men	Women
Female (0/1)	0.47	0.00	1.00
	(0.50)	(0.00)	(0.00)
Age (years)	21.48	21.41	21.56
	(3.15)	(3.11)	(3.20)
Married (0/1)	0.04	0.02	0.05
	(0.19)	(0.15)	(0.22)
Work (0/1)	0.83	0.81	0.85
	(0.38)	(0.39)	(0.36)
Post-secondary education (0/1)	0.19	0.15	0.24
	(0.39)	(0.36)	(0.42)
Children (0/1)	0.03	0.02	0.03
	(0.16)	(0.15)	(0.17)
Denmark (0/1)	0.26	0.25	0.26
	(0.44)	(0.43)	(0.44)
Norway (0/1)	0.41	0.41	0.41
	(0.49)	(0.49)	(0.49)
Sweden (0/1)	0.34	0.34	0.33
	(0.47)	(0.47)	(0.47)
Number of observations	45950	24275	21675

Notes: The table reports the mean and standard deviation (in parentheses) for different outcomes. Sample of respondents to the SILC survey from Denmark, Norway and Sweden, aged at least 18 years old.

Table A5: Share of Women in ISCO-08 Occupations

Occupation	Share of women	Skill level
Personal care workers	0.875	2
General and keyboard clerks	0.859	2
Cleaners and helpers	0.808	1
Health professionals	0.798	4
Food preparation assistants	0.780	1
Health associate professionals	0.778	3
Customer services clerks	0.753	2
Teaching professionals	0.694	4
Subsistence farmers, fishers, hunters and gatherers	0.667	2
Sales workers	0.631	2
Legal, social, cultural and related associate professionals	0.602	3
Legal, social and cultural professionals	0.579	4
Numerical and material recording clerks	0.560	2
Other clerical support workers	0.548	2
Business and administration associate professionals	0.547	3
Personal service workers	0.515	2
Business and administration professionals	0.502	4
Hospitality, retail and other services managers	0.440	3
Agricultural, forestry and fishery labourers	0.423	1
Administrative and commercial managers	0.376	4
Food processing, wood working, garment and other craft	0.373	2
Street and related sales and service workers	0.364	1
Assemblers	0.362	2
Stationary plant and machine operators	0.344	2
Production and specialised services managers	0.337	4
Handicraft and printing workers	0.307	2
Science and engineering professionals	0.306	4
Market-oriented skilled agricultural workers	0.271	2
Refuse workers and other elementary workers	0.260	1
Labourers in mining, construction, manufacturing and transport	0.225	1
Information and communications technology professionals	0.223	4
Chief executives, senior officials and legislators	0.212	4
Protective services workers	0.208	2
Information and communications technicians	0.191	3
Science and engineering associate professionals	0.173	3
Non-commissioned armed forces officers.	0.170	4
Armed forces occupations, other ranks.	0.125	1
Commissioned armed forces officers.	0.119	4
Market-oriented skilled forestry, fishery and hunting workers	0.071	2
Drivers and mobile plant operators	0.070	2
Metal, machinery and related trades workers	0.044	2
Building and related trades workers, excluding electricians	0.033	2
Electrical and electronic trades workers	0.032	2

Notes: Share of women in Denmark, Norway and Sweden who work in each of the 43 ISCO-08 occupations. Data from the SILC survey.

Table A6: Effects of Paternity Leave Reforms on Labor Force Participation and Human Capital

	Work (0/1)		Post-secondary education (0/1)	
	Men	Women	Men	Women
All countries				
RDD coef.	0.002	0.010	0.006	-0.007
(SE)	(0.011)	(0.011)	(0.009)	(0.012)
Nb. observations	24275	21675	23815	21240

Notes: The table reports RDD estimates from local polynomial regression of order 1 and corresponding to coefficient β in equation (1). Robust standard errors are reported in parentheses. Bandwidth of 20 quarters around the reform cutoff. The running variable is the quarter of birth. Sample of respondents to the SILC survey, aged 18 or above, from Denmark, Norway and Sweden.

Table A7: Robustness Checks – SILC Sample

	Female-dominated occupation (0/1)	
	Men	Women
	(1)	(2)
Main specification (Poly. = 1, BW=20)		
RDD coef.	0.045 **	0.000
(SE)	(0.018)	(0.013)
Nb. observations	13573	11643
Poly. = 2		
RDD coef.	0.044	-0.007
(SE)	(0.029)	(0.021)
Nb. observations	13573	11643
BW = CCT		
RDD coef.	0.044 **	0.000
(SE)	(0.019)	(0.013)
Nb. observations	12163	11643
BW = 14		
RDD coef.	0.049 **	0.003
(SE)	(0.022)	(0.016)
Nb. observations	9391	7935
BW = 16		
RDD coef.	0.045 **	-0.008
(SE)	(0.020)	(0.015)
Nb. observations	10703	9205
BW = 18		
RDD coef.	0.044 **	-0.004
(SE)	(0.019)	(0.014)
Nb. observations	12163	10375
BW = Honest; Poly. = 1		
RDD coef.	0.044 **	0.000
(SE)	(0.019)	(0.014)
Honest 95% CI	[0.005 ; 0.083]	[-0.030 ; 0.030]
Nb. observations	12905	9780

Notes: The table reports RDD estimates based on equation (1) and from local polynomial regression with different polynomial orders. Robust standard errors are reported in parentheses. The lower panels also report estimates with different bandwidths around the reform cutoff, including the data-driven bandwidth suggested by the procedure of Calonico et al. (2014b) and the "honest confidence intervals" from Kolesár and Rothe (2018). Sample of respondents to the SILC survey, aged 18 or above, from Denmark, Norway and Sweden.

Table A8: Robustness Checks related to Seasonality – IAT Sample

	Standardized Implicit Association Test		
	Both	Men	Women
	(1)	(2)	(3)
Panel A: Calendar month-of-birth FE			
RDD coef.	-0.035	-0.204 ***	0.047
	(0.033)	(0.066)	(0.046)
Nb. observations	3104	1080	2020
Panel B : RDD DiD			
Cutoff pre-reform	0.014	-0.016	0.033
	(0.023)	(0.046)	(0.037)
Cutoff * Reform year	-0.053 **	-0.244 **	0.032
	(0.024)	(0.105)	(0.043)
Nb. observations	5234	1839	3385

Notes: The first row "Calendar month-of-birth FE" reports RDD estimates from local polynomial regression of order 0 and controlling for calendar month-of-birth fixed effects. Bandwidth of 24 months around reform cutoff. The second row "RDD DiD" reports results from a RDD augmented with a difference in differences component that captures season of births effects. Bandwidth of 6 months around cutoff for the RDD, and 5 pre-reform years for the DiD. Standard errors (reported in parentheses) are clustered at the birth of month level. Sample of respondents who live in Belgium, Denmark, France, Iceland, Norway and Sweden.

Table A9: Robustness Checks related to Seasonality – SILC Sample

	Female-dominated occupation (0/1)	
	Men	Women
	(1)	(2)
Panel A: Calendar quarter-of-birth FE		
RDD coef.	0.046 **	-0.001
	(0.018)	(0.013)
Nb. observations	13573	11643
Panel B : RDD DiD		
Cutoff pre-reform	0.000	0.008
	(0.008)	(0.006)
Cutoff * Reform year	0.039 **	0.005
	(0.019)	(0.016)
Nb. observations	16391	14212

Notes: The first row "Calendar quarter-of-birth FE" reports RDD estimates from local polynomial regression of order 1 and controlling for calendar quarter-of-birth fixed effects. Bandwidth of 20 quarters around reform cutoff. The second row "RDD DiD" reports results from a RDD augmented with a difference in differences component that captures season of births effects. Bandwidth of 2 quarters around cutoff for the RDD, and 6 years for the DiD. Robust standard errors are reported in parentheses. Sample of respondents to the SILC survey, aged 18 or above, from Denmark, Norway and Sweden.

Table A10: Survey of the Literature on the Effects of Paternity Leave Reforms on Fathers' Time Dedicated to Childcare / Housework

Belgium	**Fontenay and Tojerow (2024)*** reveal that 10 years after the birth of their first child, the fathers eligible for paternity leave after the 2002 reform in Belgium spent on average 45 more minutes per day taking care of their children, that is twice as much as the time spent by ineligible fathers.
Denmark	**Huerta et al. (2013)†** observe that the fathers who take leave in Denmark are more involved with their 6-month-old child (including feeding, changing diapers, getting child to bed), compared to fathers who do not take leave.
France	**Pailhé et al. (2024)*** find that the households where the father took a two-week paternity leave are dividing more equally all child-rearing tasks (bathing, changing, putting children to bed and night caring) two months after childbirth. The effect is stronger among first-time parents.
Iceland	**Arnalds et al. (2013)*** compare first-time parents whose child was born in 1997 (pre-reform) and 2003/2009 (post-reform) and look at the division of child-rearing tasks within the household up to three years. Post-reform children were between 13 and 23 percentage points more likely to receive care from both parents.
Norway	**Kotsadam and Finseraas (2011)*** find that the parents who had their last born child right after the 1993 daddy-quota reform in Norway are 50 percent more likely to equally divide domestic tasks (e.g. washing clothes). The authors also find a reduction in conflicts over household division of labor.
Sweden	**Haas and Hwang (2008)†** observe that fathers in large private companies in Sweden who take longer than average leave are more involved in childcare-related tasks.
Other countries	**Farré and González (2019)*** find that fathers in Spain who have been eligible for paternity leave in 2007 did almost an hour more childcare per day in 2009–10 compared with ineligible fathers. **Patnaik (2019)*** evaluates the effect of the Quebec Parental Insurance Plan in Canada and reveals that exposed fathers spent more time physically at home and dedicated more time to housework per day. **Tamm (2019)*** find that in Germany, the fathers who took a paternity leave in 2007 dedicate 1.2 more hours to their family per weekday and an additional 1.4 more hours to childcare on Saturdays and 1.6 hours on Sundays. They also increase the time they dedicate to housework (washing, cooking, cleaning) by 0.5 hours per weekday.

* Causal identification strategy, † correlational evidence