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# Fathers but not caregivers*

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*Fathers' parental leave uptake remains low in many advanced economies despite substantial policy efforts. We study a setting where financial and eligibility barriers are minimal: employed, native-born first-time fathers entitled to generous, non-transferable leave benefits. Using Swedish population register data for 1995–2015, we document three key facts: (i) low uptake follows a persistent U-shaped income gradient, (ii) its determinants vary across the distribution—economic constraints at the bottom and top, workplace norms in the middle—and (iii) these constraints have grown more salient over time. Quota reforms increased uptake on average but did not narrow differences between constrained and unconstrained fathers. Using quasi-random sibling-sex composition, we show that exposure to traditional gender-role environments increases the likelihood of low uptake, especially in recent cohorts. The results highlight the limits of financial incentives and point to workplace and household norms as central barriers to equal parental leave participation.*

**Keywords**: Men; parental leave; gender norms; father's quota

**JEL Classification**: D13; J13; J16; J18

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## 1. Introduction

In recent decades, an increase in fathers' use of parental leave has been identified as a crucial step toward greater economic gender equality (Angelov et al., 2016; Kleven et al., 2019; Canaan et al., 2022). To promote a more equitable sharing of childcare responsibilities, many countries have implemented policies reserving a portion of paid parental leave for fathers.[1] These quotas have raised fathers' uptake, yet a substantial share still forgoes their entitlement, even when the leave is non-transferable and generously compensated (OECD, 2016; Koslowski, 2022).

This puzzle is particularly striking in Sweden, a long-standing leader in gender-equal family policy, where fathers have had equal formal rights to parental leave for over fifty years. The Swedish system combines nearly universal eligibility with a high earnings replacement rate—around 80 percent up to a generous ceiling—minimizing financial disincentives to take leave. Yet only about half of fathers use their reserved quota. Moreover, uptake is not randomly distributed: as shown in Figure 1, low uptake follows a U-shaped pattern across the income distribution, with the highest rates of non-use among both low- and high-income fathers. This pattern is inconsistent with a single economic mechanism and instead suggests heterogeneity in factors that limit fathers' leave-taking across the earnings distribution. Building on evidence that fathers respond strongly to financial incentives when using earmarked leave (Jørgensen & Søgaard, 2024), Sweden's high-replacement-rate-system offers an ideal setting to examine what constrains uptake once these incentives are less binding.

We study the factors underlying persistently low uptake and how they differ across the income distribution, using population-wide administrative data on Swedish fathers. Fathers' leave-taking may be limited by physical barriers such as poor health, incarceration, or separation; by economic circumstances including self-employment, unstable work, or a high household income share; and by social norms that discourage paternal leave use. To capture the

[1] In 2019, the European Parliament mandated all members to reserve at least two months of paid parental leave for each parent (Directive 2019/1158).

normative dimension, we combine measures capturing influences at different levels: workplace environments dominated by men with low leave uptake, regional gender attitudes, and early-life exposure to gender roles (proxied by sibling-sex composition). The analysis focuses on first-time, Swedish-born fathers who were employed in the year before childbirth and cohabiting with the child's mother, ensuring that eligibility and financial constraints are minimal. We define low uptake as taking at most half of the individually reserved days during the child's first two years of life, a period critical for fostering lasting involvement in childcare (Cools et al., 2015).

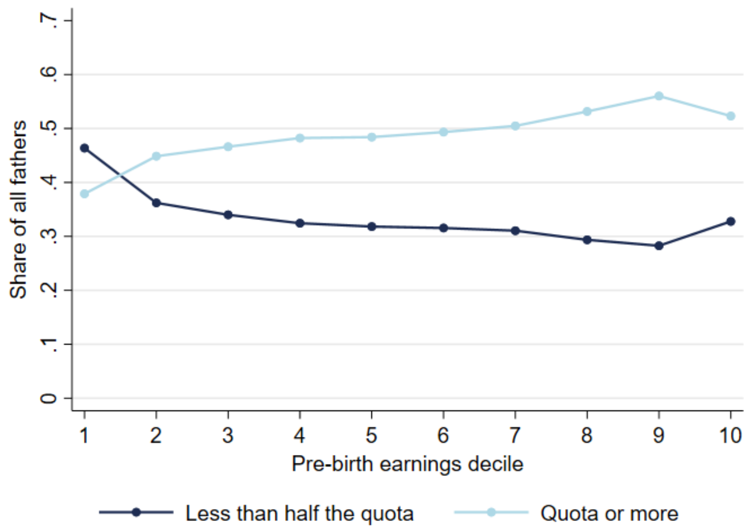


**Figure 1. The share of all fathers by average degree of parental leave uptake for children born 1995–2015 and earnings decile**
*Notes:* The figure shows the share of fathers of children born in 1995–2015, by parental leave uptake and pre-birth earning decile. Parental leave uptake is measured as the number of paid parental leave days during the child's first two years. The light blue line indicates the share of fathers taking at least their full reserved quota (30 days from 1995 to 2001 and 60 days from 2002). The black line indicates the share taking at most half of the reserved quota (15 days in 1995–2001 and 30 days from 2002 onward). Pre-birth earnings are measured in the calendar year before childbirth and include annual labor income from wage- and self-employment as well as capital income.

Our empirical analysis integrates descriptive and causal evidence to understand why fathers' parental leave uptake remains low. We begin by mapping fathers' parental leave uptake across the earnings distribution and over time, establishing three key empirical facts: (i) There is persistent U-shaped relationship between pre-birth income and low uptake, which has flattened but not disappeared over time: it has declined mainly among middle- and high-income fathers, while stagnating at the bottom of the distribution; (ii) The factors associated with low uptake differ systematically across the income distribution: physical constraints are rare, workplace norms (our proximate measure of social constraints) are most influential among middle-income fathers, and economic constraints dominate at the bottom and the top; (iii) . In all income groups, observable constraints have become increasingly salient, but the underlying mechanisms vary. Workplace norms have grown in relevance for middle-income fathers, while economic opportunity costs have become more important for high-income fathers.

We then turn to test whether institutional design can offset these barriers, by assessing the causal effects of two reforms—the 1995 and 2002 expansions of earmarked leave—using a difference-in-discontinuity design. These reforms substantially increased the number of reserved paid parental leave days for fathers and were intended to promote more equal sharing of childcare. By comparing fathers just before and after each reform, we test whether expanding individual quotas primarily raises uptake among those already likely to take leave, or whether it also alleviates the economic and normative barriers identified above. The results show that while the reforms increased overall uptake, they did not differentially raise participation among constrained fathers. This suggests that while constraints did not prevent fathers from responding to the reforms, institutional changes alone are insufficient to overcome the barriers identified in the descriptive analysis.

Finally, to understand why low uptake persists despite generous compensation and repeated policy reforms, we examine whether gender role attitudes shape fathers' leave-taking behavior. Because such norms are inherently difficult to observe directly, we rely on two proxies to capture different dimensions of normative influence. We show that fathers residing in regions with more traditional gender attitudes, measured using data from the World Values Survey, are significantly more likely to forgo leave. In addition, using quasi-experimental evidence from sibling-sex composition (Brenøe, 2021) as a proxy for early-life exposure to gender-role environments, we find that fathers from more gender-traditional family environments are also less likely to take leave. Specifically, having a younger sister rather than a younger brother increases the probability of low uptake by about 1 percentage point, corresponding to roughly a 4 percent increase relative to the mean. While economically modest, this magnitude is comparable to effect sizes in Brenøe (2021). This effect is concentrated among recent cohorts, consistent with an increasingly important role of early-life gender norms in shaping fathers' caregiving behavior.

Prior research has shown that paternity leave reforms raise fathers' average uptake and has studied in detail the intensive margin, i.e., how leave is divided between mothers and fathers (e.g., Sundström & Duvander, 2002; Duvander & Johansson, 2012; Ekberg et al., 2013; Dahl et al., 2014; Patnaik, 2019). Much less is known about why some fathers continue to abstain from leave altogether, even when entitlements are generous. The few existing studies describe the characteristics of these fathers. Ma et al. (2020) highlight that young, low-income, and foreign-born fathers are less likely to take leave, while Fahlén & Duvander (2021) and Saarikallio-Torp & Miettinen (2021) point to relationship status, education, and job characteristics. However, these studies remain largely descriptive and do not disentangle whether low uptake reflects economic constraints, preferences, or persistent social norms.

This paper makes four contributions. First, it shifts attention from the intensive to the extensive margin in a context where the average uptake of parental leave is high, focusing on fathers with minimal eligibility and financial barriers who nonetheless abstain, thereby isolating the role of non-financial barriers. Second, it provides new descriptive evidence on the determinants of low uptake, documenting a persistent U-shaped income gradient, systematic variation in the types of constraints across the distribution, and changes in their relative importance over time. Third, it offers causal evidence from two sources—policy reforms expanding earmarked parental leave (1995 and 2002) and quasi-experimental variation in early-life exposure to gender role environments—showing how institutional design and social norms shape behavior. Fourth, it contributes to broader debates on the interaction between policy and culture (Jørgensen & Søgaard, 2024; Albrecht et al., 2024; Andresen & Nix, 2024), demonstrating that while reforms raised average parental leave uptake among fathers, persistent gender norms continue to constrain uptake and have become increasingly salient over time.

The paper proceeds as follows. Section 2 describes the data. Section 3 presents descriptive evidence and the three stylized facts. Section 4 analyzes the quota reforms using a difference-in-discontinuity design. Section 5 examines the role of gender norms at regional and individual levels. Section 6 concludes.

## 2. Sample and constraints

### *2.1 Data sources*

We utilize data on parental leave from the MiDas database at the Swedish Social Insurance Office, which covers all parental leave benefit payments since 1994.[2] We link these records to administrative registers from Statistics Sweden containing information on socioeconomic characteristics, such as age, education, employment, workplace, earnings, and family structure. We obtain health data from the inpatient care register at the Social Insurance Office and information on criminal convictions from the official crime register compiled by the National Council for Crime Prevention. Finally, we use data on regional gender attitudes from the World Values Survey (WVS).

[2] The data includes details about the type of benefit, net and gross days, amounts, and the dates the benefit covers. This information is reported by the child and beneficiary. We use the net days of parental leave benefits, meaning that days with partial replacement are combined so that one day equals full-time replacement. Duvander & Viklund (2014) show that the correlation between leave days and benefits used is very high for fathers, implying that unpaid parental leave should not be a problem in our case.

### 2.2 Sample restrictions

The main analytic sample consists of first-time fathers whose children were born in Sweden between 1995 and 2015. We focus on first-time fathers because all observed characteristics can be measured before the birth of the first child, avoiding confounding from prior parenting experience or established leave-taking patterns.

To focus on fathers for whom the short-term financial cost of taking leave is low, we limit the sample to those who are employed and had a registered workplace in the year before childbirth. This aligns with the stated aim of the Swedish parental leave system, which is designed to enable employed individuals to stay at home with their children. Employed parents have been entitled to at least 77.6 percent wage replacement throughout the period. Many employees are also covered by collective insurances that increase the replacement pay during parental leave even further (Sjögren Lindquist and Wadensjö, 2005).

For consistency, we exclude fathers who were full-time students in the year before childbirth, as they are likely to have little or no earnings and face distinct constraints. To ensure comparability in caregiving arrangements, we also restrict the sample to fathers who were cohabiting with the child's mother in the year of birth or the following year. Non-cohabiting couples are likely to face different institutional incentives and caregiving norms and differ in observable characteristics. Finally, to ensure complete data and consistent exposure to the Swedish parental leave system, we restrict the analysis to Swedish-born fathers and children born in Sweden. This improves data coverage, particularly for variables related to family background such as sibling composition and parental education, which are central to our analysis of social norms.

The restrictions yield a more homogeneous and policy-relevant sample of employed, Swedish-born first-time fathers with relatively few structural barriers to leave-taking. As shown in Web Appendix W2, the analytic sample closely resembles the full population of fathers, with the main difference being that the full population also includes non-employed fathers.

### *2.3 Measures of low uptake and constraints*

To capture genuinely low engagement in parental leave, we define low uptake as using at most half of the reserved days with paid parental leave within the first two years after birth.[3] Fathers below this threshold are not merely falling slightly short but are well below the policy's intention of promoting equal caregiving. As such, this measure identifies those who make limited use of their entitlement and are unlikely to contribute substantially to early childcare. The cutoff varies with policy reforms that extended the number of reserved days: 0–15 days for children born 1995–2001 and 0–30 days for children born 2002–2015. We focus on the two-year window (as in Duvander & Johansson, 2019) for two reasons. First, leave taken during this period is most closely associated with active caregiving before preschool and aligns more with the reform's goal of promoting early father involvement (e.g., Cools et al, 2015). Second, the two-year horizon facilitates cross-country comparisons with less flexible systems in other countries (OECD, 2021). We exclude periods of double days (when both parents are on leave at the same time), as caregiving is shared, and these days do not count toward the individual quota (see Web Appendix W1 for institutional details).

Income deciles are constructed from fathers' pre-birth annual earnings, measured in the calendar year before childbirth. Earnings include wage income, self-employment income, and capital income. Deciles are defined each year using all fathers in the sample, including those with zero income.

[3] Since 2002, parents have shared 480 days of leave (450 days prior to 1995), of which 390 are income-based and 90 are paid at a low flat rate. We do not distinguish between these types. We do however exclude the ten days of birth-related paternity leave to focus on sustained caregiving rather than the near-universal short leave taken immediately after birth.

We group potential barriers to fathers' parental leave uptake into three broad categories: physical, economic, and social norms. We operationalize these using a combination of administrative and survey-based measures that capture the main channels through which caregiving capacity, eligibility, opportunity costs, and cultural expectations can limit fathers' parental leave uptake. While physical constraints reflect direct limits on fathers' ability to provide care, economic and social-normative constraints operate through incentives and behavioral responses.

We capture physical constraints on parental leave uptake through indicators of health, criminal history, and living arrangements. A father is classified as being in poor health if he was hospitalized for at least seven consecutive days in the year before birth or received sickness benefits, which are granted after more than two weeks of illness-related absence from work. Poor health may reduce parenting capacity and self-efficacy, potentially discouraging leave-taking (Angst & Deatrick, 1996). To proxy incarceration, we use an indicator for whether the father was convicted of a non-traffic crime at any time before the child's birth year, acknowledging that prison sentence data are not available. Incarceration presents a clear barrier to caregiving, and fathers with criminal records may be less available or welcomed in caregiving roles due to stigma or family stress (Roettger & Swisher, 2013; Dobbie et al., 2019).[4] Finally, we identify fathers who were no longer cohabiting with the child's mother in the second year after childbirth as not living full-time with the child, a barrier we refer to as separation. While shared or sole custody may still allow leave-taking, non-custodial arrangements typically inhibit it. In Sweden, separation has been shown to reduce fathers' parental leave uptake (Fahlén & Duvander, 2021).

[4] There is a literature of parenthood as a potential "turning point" in a criminal career, see e.g., Monsbakken et al. (2013). Our measure is meant to capture the possible effects of criminal activity before the child is born on parental leave uptake.

To capture economic constraints, we use proxies for high opportunity costs or household bargaining power: unstable work, self-employment, and high household income share. A father is classified as having unstable work if he either changed employer or experienced non-employment in the two years before childbirth, indicated by a missing employer ID or receipt of unemployment benefits. Employment instability may signal job insecurity, which can deter leave-taking due to fear of income loss or career disruption (Sundström & Duvander, 2002). A father is classified as self-employed if the majority of his labor income derives from self-employment, whether through incorporated or unincorporated firms. Self-employed fathers may also face high opportunity costs of absence, not only due to lost income but also because their leave could threaten the viability of their business. Lastly, we define a father as facing an economic constraint if he contributed more than 70 percent of household income in the year before childbirth. This threshold proxies high relative earnings and household dependence, which may increase the opportunity cost of leave and shift intra-household bargaining power. Both unitary and collective models of household decision-making predict that parents with greater economic leverage, typically high-earning fathers, are less likely to take leave (Becker, 1965; Lundberg & Pollak, 1996; Manser & Brown, 1980).

We recognize that social norms are inherently difficult to observe directly and therefore rely on multiple proxies at the workplace, regional, and individual levels to capture different dimensions of normative influence. At the workplace level, we use register data to construct an indicator for whether the father is employed at a workplace where (i) at least 80 percent of employees are men and (ii) the average parental leave uptake among male colleagues with children born in the preceding two years is at most half of the reserved days. This measure provides a granular, proximate indicator of the social environment fathers face and captures exposure to male-dominated, low-uptake workplaces that may reinforce traditional expectations around gender roles and breadwinning.[5] Prior research documents strong peer effects in paternal leave uptake, with fathers being less likely to take leave when few of their colleagues do so (Bygren & Duvander, 2006; Dahl et al., 2014; Carlsson & Reshid, 2022; Tallås Ahlzén, 2022; Casarico et al, 2025). Because workplace sorting may reflect occupational choices, we treat this measure as descriptive rather than causal.

At the regional level, we use data from the World Values Survey (WVS) to construct an index of traditional gender attitudes. Specifically, we calculate the share of respondents in a father's region who agree with at least one of the following statements: "*A university education is more important for a boy than for a girl*", "*Men make better political leaders than women*", or "*When jobs are scarce, men should have more right to a job than women*". These statements capture beliefs about gender roles in public and private life and are widely used to proxy gender-ideological contexts (e.g., Gornick, 2015; Bloksgaard, 2015). While regional attitudes may also reflect family sorting, they capture broader cultural environments that extend beyond individual workplaces.

At the individual level, we use the father's sibling sex composition as a proxy for early-life exposure to gender norms (see Section 7.2 for more details). Specifically, we use an indicator for whether the father has a younger sister, building on findings that growing up in mixed-gender sibling groups is associated with more traditional gendered behaviors and occupational choices (Brenøe, 2021). This proxy is plausibly quasi-random and provides the most exogenous variation in normative exposure.

Table 1 summarizes the prevalence of each type of constraint across the pre-birth earnings distribution for all fathers and for fathers with a low uptake of parental leave benefits. For ease of presentation, we group fathers into three earnings categories: the bottom decile (decile 1),

[5] This measure may partly capture workplace characteristics rather than norms. However, by combining male dominance with persistently low uptake, it is more likely to reflect prevailing social norms and peer effects around fatherhood and breadwinning than structural job constraints alone.

the middle (deciles 2–9), and the top decile (decile 10), where patterns of leave uptake are relatively homogeneous (see Figure 1). Additional justification for this grouping is provided in Web Appendix Table W1.

**Table 1. Prevalence of constraints among all fathers and among fathers with a low uptake, by fathers' pre-birth earnings decile**

| | All | | | Low uptake | | |
|---|---|---|---|---|---|---|
| | Decile 1 | Deciles 2–9 | Decile 10 | Decile 1 | Deciles 2–9 | Decile 10 |
| **Economic constraints** | | | | | | |
| Changed job | 0.235 | 0.171 | 0.236 | 0.221 | 0.162 | 0.230 |
| Previously unemployed | 0.342 | 0.127 | 0.022 | 0.334 | 0.153 | 0.026 |
| Self-employed | 0.213 | 0.044 | 0.066 | 0.270 | 0.066 | 0.096 |
| High share of household income | 0.147 | 0.128 | 0.344 | 0.145 | 0.144 | 0.452 |
| Share with any economic constraints | 0.767 | 0.416 | 0.544 | 0.786 | 0.459 | 0.633 |
| Average number of economic constraints | 0.937 | 0.470 | 0.668 | 0.969 | 0.525 | 0.804 |
| **Physical constraints** | | | | | | |
| Hospitalized | 0.003 | 0.002 | 0.001 | 0.003 | 0.002 | 0.002 |
| Sick leave | 0.091 | 0.056 | 0.020 | 0.090 | 0.061 | 0.021 |
| Crime | 0.115 | 0.055 | 0.033 | 0.123 | 0.066 | 0.043 |
| Separation | 0.106 | 0.044 | 0.030 | 0.121 | 0.059 | 0.044 |
| Share with any physical constraints | 0.265 | 0.142 | 0.080 | 0.284 | 0.168 | 0.104 |
| Average number of physical constraints | 0.314 | 0.157 | 0.084 | 0.338 | 0.189 | 0.110 |
| **Norm constraints** | | | | | | |
| 80% men and low PL uptake at workplace | 0.529 | 0.376 | 0.221 | 0.595 | 0.473 | 0.288 |
| **Share with any constraint** | 0.898 | 0.656 | 0.641 | 0.921 | 0.735 | 0.740 |
| **Average number of constraints** | 1.780 | 1.003 | 0.973 | 1.902 | 1.187 | 1.201 |
| Observations | 53,098 | 424,702 | 53,075 | 24,623 | 135,249 | 17,398 |

*Notes:* The table reports the prevalence of economic, physical, and norm-related constraints for first-time fathers of children born in 1995–2015, by fathers' pre-birth earnings decile. Entries show shares of fathers with each characteristic unless otherwise noted. Deciles refer to fathers' pre-birth earnings. Pre-birth earnings are measured in the calendar year before childbirth and include annual labor income from wage- and self-employment as well as capital income.

Economic constraints are concentrated among low-income fathers. For instance, among fathers with a low uptake 79 percent in decile 1 face at least one economic constraint, compared to 46 percent in deciles 2–9 and 63 percent in decile 10. Physical constraints are also more common among low earners, driven by higher rates of criminal convictions and separation. Normative constraints, proxied by exposure to traditional gender norms at the workplace, decline steadily with earnings (see Section 7 for descriptive statistics on gender norms at the regional and individual levels).

Across all earnings groups, fathers with low leave uptake are systematically more constrained than fathers overall. They are more likely to face at least one constraint and are exposed to a greater number of economic, physical, and normative constraints within every earnings group, with particularly pronounced differences in the middle and top deciles. Among all fathers, 90 percent in the lowest decile face at least one constraint, compared to 64 percent in the top decile and 66 percent in the middle. A similar pattern, but at consistently higher levels, is present among fathers with low uptake: 92 percent of fathers in the lowest decile face at least one constraint, compared to 74 percent in the top decile and 73 percent in the middle. Web Appendix W2 shows that the prevalence of these constraints has remained stable over time.

## 3. Mapping constraints on fathers' parental leave uptake

We first document how physical, economic, and social constraints are associated with fathers' low uptake of parental leave across the income distribution and how their relative importance has evolved over time. We quantify these descriptive patterns using linear probability models of the form:

$$PL_{it} = \alpha + \gamma Income_d + \tau Constr_{it} Income_d + \omega Constr_{it} + \varepsilon_{it} \quad (1)$$

where $PL_{it}$ is one if the father *i,* with a child born in year *t*, exhibits low uptake, defined as taking no more than half of the reserved quota days, and zero otherwise. The model includes the father's earnings decile in the year before childbirth, $Income_d$, and an indicator for whether the father is subject to a physical, economic, or social constraint, $Constr_{it}$. The interaction term allows the association between constraints and uptake to vary across the earnings distribution. In this descriptive analysis, social constraints are focused at the workplace level because it is our most proximate and directly observable measure of normative influence.

For time trends, we estimate a similar specification with yearly dummies:

$$PL_{it} = \alpha + \gamma Year_t + \tau Constr_{it} Year_t + \omega Constr_{it} + \varepsilon_{it} \qquad (2),$$

where $Year_t$ is a yearly dummy variable. We estimate this model separately for three earnings groups—bottom decile, middle deciles (2–9), and top decile—given their distinct patterns of low uptake.

In both specifications, we plot sum of intercept ($\alpha$) and the relevant coefficients ($\gamma$), either unconditionally (black line) or conditional on constraints (blue and grey lines). The gap between unconditional and conditional probabilities provides a descriptive measure of the importance of how strongly each type of constraint is associated with low parental leave uptake. The descriptive evidence reveals three clear patterns: (1) a persistent U-shaped income gradient in low uptake; (2) systematic differences of constraints across the income distribution; and (3) a diverging pattern of low uptake over time.

*3.1 Low uptake follows a persistent U-shaped income pattern.*

Figure 2, Panel A, presents estimates using equation (1). As observed in Figure 1, fathers at both the bottom and the top of the income distribution are substantially more likely to take no more than half of the reserved quota days, while middle-income fathers are least likely to do so. This U-shaped pattern persists across cohorts, although overall uptake has increased steadily over time (see Figure W2 in Web Appendix W1).

*3.2 Constraints vary over the income distribution*

Conditioning on having any constraint— physical, economic, or social (dashed grey line in Figure 2 , Panel A)—is associated with a reduction in the likelihood of low uptake by about 25 percentage points among fathers in deciles 2–10, and about 36 percentage points in the bottom decile. This corresponds to, on average, a reduction of 20–25 percent relative to the unconditional probability. Importantly, while the overall importance of having any constraint is substantial throughout the income distribution, the relative importance of specific constraints varies across income groups.

Physical constraints are rare and have negligible importance (dotted blue line), suggesting that illness, incarceration, or separation are not major determinants of low leave uptake for this group of fathers. By contrast, workplace norm constraints (solid blue line) emerge as the most influential factor in explaining low parental leave uptake, particularly among fathers in the middle of the income distribution, where they are associated with a 15 percent reduction in low uptake relative to the unconditional probability. Economic constraints (blue dashed line in Figure 2, Panel A (and disaggregated in Panel B) dominate at the top of the distribution, but appear to be important across the entire distribution: conditioning on economic constraints decreases low uptake by about 20 percent among top earners, 9 percent among low earners, and 7 percent among middle-income fathers. However, the mechanisms differ. Figure 2, Panel B, shows that being the primary household earner is the main barrier for top earners, while self-employment is most important at the bottom of the distribution—both factors consistent with higher opportunity costs. Middle-income fathers are least affected by economic constraints overall, though self-employment remains relevant for them as well.

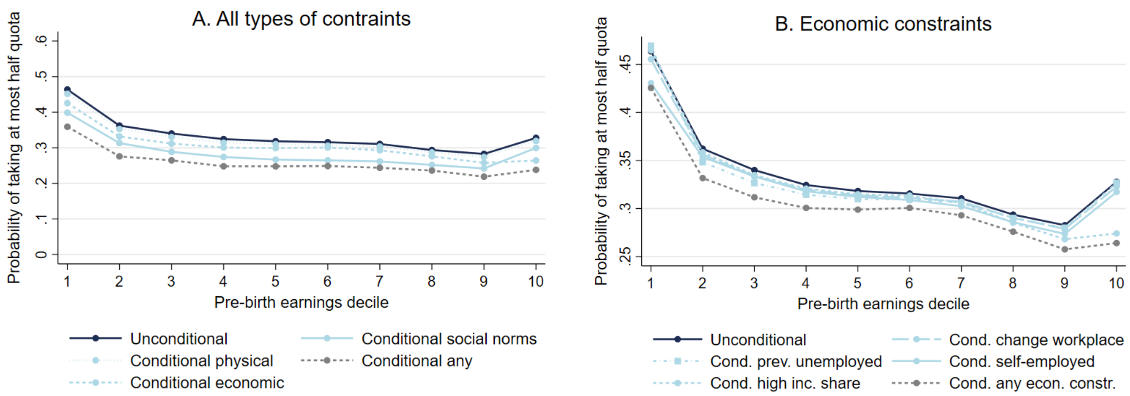


**Figure 2. Probability of low parental leave uptake by earnings decile, unconditional and conditional on constraints**

*Note:* The figure shows the probability that fathers of children born in 1995–2015 take at most half of their reserved parental leave quota, by pre-birth earnings decile. Panel A (left-hand side) plots unconditional estimates (black solid line) and conditional estimates controlling for physical (blue dotted), economic (blue dashed), and workplace norm constraints (blue solid). The grey dashed line shows estimates conditional on any constraint. Panel B (right-hand side) disaggregates economic constraints, plotting conditional estimates for fathers who changed workplace (blue dashed), were previously unemployed (blue dotted), were self-employed (blue solid), or contributed a high share of household income (blue short-dashed). The grey dashed line shows the joint contribution of any economic constraint. Half of the quota corresponds to 0–15 days for children born in 1995–2001 and 0–30 days for those born from 2002 onward.

Overall, the effects of different constraints are largely additive at the aggregate level. As shown in Figure 2, Panel A, the dashed grey line, representing any type of constraint, is consistently below the lines for each individual constraint, indicating that the combined explanatory power of multiple barriers exceeds that of any single one. A similar pattern appears within the set of economic sub-constraints (Figure 2, Panel B): each factor explains only a modest share of the variation in low uptake, while conditioning on all of them together yields a much larger reduction.

*3.3 Diverging patterns of low uptake over time*

Figure 3 presents trends in low parental leave uptake across the three income groups. Between 1995 and 2015, uptake improved for middle- and high-income fathers but stagnated among low-income fathers, who consistently exhibit the highest rates of low uptake. Specifically, the share of low-income fathers taking less than half of the quota declined only modestly, from 45 to 39 percent, compared with a decline from 33 to 25 percent among fathers in deciles 2–9 and from 37 to 24 percent in the top decile.

At the same time, the gap between unconditional and conditional uptake widened for all groups, suggesting that observable constraints have become increasingly important. In decile 1, this gap grew from 15.8 percent in 1995 to 23.3 percent (relative to unconditional uptake). In the middle deciles, the corresponding increase was from 16.9 to 32.1 percent, and in the top decile, from 14.4 to 30.3 percent.

The underlying mechanisms, however, again diverge across income groups. Among middle-income fathers, workplace norms have become increasingly important, rising from 9.4 percent in 1995 to 22.4 percent in 2015. Economic constraints show a similar but less pronounced trend. Meanwhile, the trend in economic constraints is especially stark in the top decile. In particular, we observe a growing importance of earning a high share of the household income among these fathers. The results indicate that opportunity costs linked to being the primary earner play a central role in leave decisions among high-income fathers, although norms play an increasing importance also here. For low-income fathers, the role of individual constraints appears more diffuse. Conditional probabilities show no clear pattern by constraint type, but the growing gap between conditional and unconditional uptake suggests that these fathers increasingly face multiple constraints rather than a single dominant barrier.

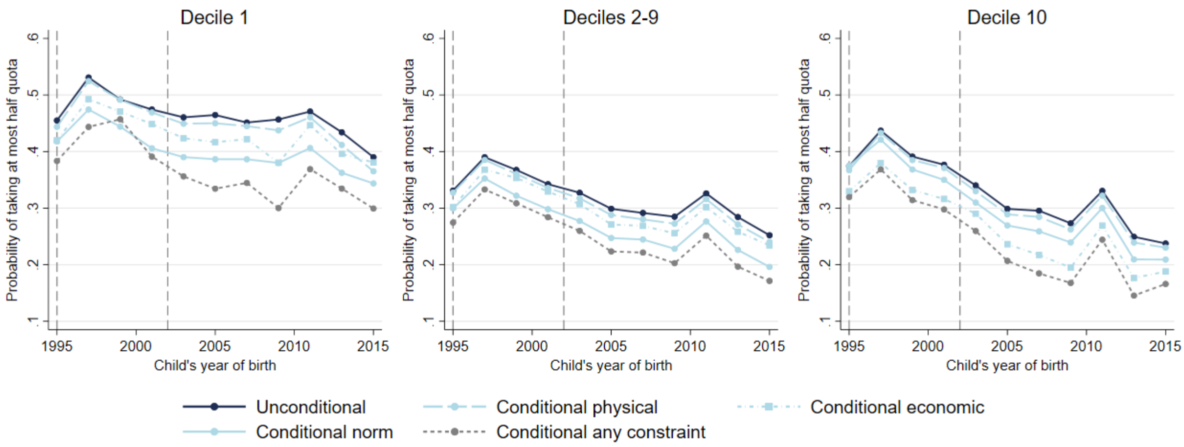


**Figure 3. The share of fathers by degree of parental leave uptake in pre-birth income decile by children's year of birth**

Note: The figure shows the share of fathers who take at most half the quota of the parental leave by birthyear of the child unconditionally on any constraint (black solid line), conditional on any physical constraint (blue dotted line), conditional on any economic constraint (blue dashed line), conditional on social norms at the workplace level constraint (blue solid line), and conditional on any constraint, be it physical, economic, or norms (grey dashed line). Half of the quota refers to 0-15 days from 1995 to 2001 and 0-30 days from 2002. The vertical dashed lines in 1995 and 2002 indicate the implementation of the first and second parental leave quotas. Records of parental leave benefits are unavailable for the fall of 2013, resulting in a lower amount of leave registered for fathers of children born between 2011 and 2013.

Taken together, these patterns show that the determinants of low uptake differ systematically across the income distribution and have diverged over time, underscoring the need to assess whether changes in institutional design, through the 1995 and 2002 reforms of earmarked leave, have mitigated these barriers.

## 4. The response of constrained fathers to reserved parental days

We leverage the staggered implementation of Sweden's parental leave quotas in January 1995 and January 2002 to causally assess how constrained and unconstrained fathers responded to these policy reforms. The quotas were designed to increase fathers' uptake by reserving leave days exclusively for each parent.[6] As shown in Figure 3, the relevance of constraints varies over time and across income groups, suggesting that the effects of policy design may differ across fathers' circumstances. This is also consistent with prior evaluations finding that these reforms had heterogeneous effects across the income distribution (Swedish Social Insurance Agency, 2019).

Our central question in this section is whether the 1995 and 2002 reforms relaxed the barriers identified in the previous section. Both reforms expanded fathers' reserved paid parental leave days and were explicitly designed to increase uptake. Given the findings in Figure 3, that low uptake declined mainly among middle- and high-income fathers but stagnated at the bottom of the distribution, we test whether these policy changes contributed to that divergence. Specifically, we assess whether the reforms reduced low uptake more among fathers facing observable constraints—physical, economic, or normative—than among those without such barriers.

We estimate the differential reform effects using a difference-in-discontinuity (Diff-in-Disc) design that compares changes in leave uptake around each reform cutoff (January 1, in 1995 and 2002 respectively), using the pre-year as a placebo difference to absorb seasonal variation. We allow the reform effect to vary between constrained and unconstrained fathers, classifying a father as constrained if they face at least one physical, economic, or social constraint. We focus on this aggregate measure of constraints rather than disaggregated effects, for three reasons. First, some subcomponents (e.g., illness, separation) are rare, limiting statistical power. Second, aggregation provides a cleaner link to the policy question, since quotas may influence multiple constraints simultaneously, and the relevant counterfactual is whether constrained fathers overall respond differently than unconstrained fathers. Third,

[6] Further institutional details are provided in Web Appendix W1.

focusing on the aggregate avoids the multiple-testing concerns that arise when estimating many subgroup effects. We estimate the following specification:

$$PL_i = \alpha_0 + \gamma_1(Treated_i * Reform_i * Constr_i) + \gamma_2(Treated_i * Reform_i) +$$
$$\gamma_3(Treated_i * Constr_i) + \gamma_4(Reform_i * Constr_i) + \gamma_5\ Reform_i + \gamma_6 Treated_i +$$
$$\gamma_7 Constr_i + f(Birthdate_i) \times [\gamma_8 + \gamma_9 Reform_i +$$
$$\gamma_{10} Constr_i + \gamma_{11}(Reform_i * \ Constr_i)] + \varepsilon_i \qquad (3)$$

The outcome variable, $PL_i$, is an indicator equal to one if father $i$ takes less than half of the reserved parental leave days in the first two years since the child was born, i.e., low uptake. Compliance with the policy implies a negative reform effect. $Treated_i$ is an indicator variable of the child being born in the first six months of the year. $Reform_i$ equals one for fathers whose child is born within six months of the reform, and zero for placebo births (those born between 18 and 6 months before the reform). $Constr_i$ indicates if the father is constrained, including as before physical constraints, economic constraints, and social constraints. $f(Birthdate_i)$ is a second-order polynomial with triangular weights, allowed to differ on each side of the cutoff. We use robust standard errors.

The coefficient of interest, $\gamma_1$, captures the differential reform effect for constrained fathers. A negative estimate implies that reforms disproportionately reduced low uptake among constrained fathers, while a positive estimate implies weaker responsiveness. The average reform effect across all fathers is given by $\gamma_2$. The remaining coefficients serve as controls: γ3–γ7 account for baseline differences by constraint status and birth timing, while γ8–γ11 flexibly capture trends in the running variable around the cutoff.

The sample includes fathers whose child is born within 6 months of each reform. By incorporating placebo years, our identification relies on the assumption that parents cannot manipulate birth timing around the reform cutoffs and that any confounding time trends are similar across years, ensuring comparability between treated and control groups (Grembi et al., 2016). Given that precise birth dates are difficult to fully manipulate, focusing on births near the cutoff helps mitigate concerns about endogenous timing.

**Table 2. Reform analysis on low parental leave uptake using regression discontinuity**

| | (1) | (2) | (3) | (4) | (5) | (6) |
|---|---|---|---|---|---|---|
| | | Reform 1995 | | | Reform 2002 | |
| | Decile 1 | Decile 2-9 | Decile 10 | Decile 1 | Decile 2-9 | Decile 10 |
| TreatedXReformXConstr. | 0.374 | 0.041 | -0.102 | 0.298 | -0.055 | 0.022 |
| | (0.303) | (0.069) | (0.204) | (0.332) | (0.066) | (0.188) |
| TreatedXReform | -0.514* | -0.103* | -0.072 | -0.387 | 0.100* | 0.036 |
| | (0.284) | (0.056) | (0.166) | (0.317) | (0.053) | (0.147) |
| TreatedXConstr. | -0.352* | 0.035 | 0.103 | -0.133 | 0.030 | 0.088 |
| | (0.188) | (0.049) | (0.142) | (0.245) | (0.048) | (0.120) |
| ReformXConstr. | -0.115 | -0.027 | 0.053 | -0.215 | 0.034 | -0.190 |
| | (0.208) | (0.052) | (0.159) | (0.263) | (0.049) | (0.140) |
| Constrained | 0.264* | 0.070* | 0.031 | -0.006 | 0.036 | 0.208** |
| | (0.137) | (0.037) | (0.111) | (0.192) | (0.036) | (0.092) |
| Treated | 0.420** | -0.070* | -0.251** | 0.254 | -0.048 | -0.020 |
| | (0.173) | (0.040) | (0.115) | (0.234) | (0.039) | (0.085) |
| Reform | 0.087 | 0.003 | 0.062 | 0.242 | -0.061 | 0.232** |
| | (0.194) | (0.043) | (0.132) | (0.253) | (0.039) | (0.111) |
| Constant | 0.351*** | 0.338*** | 0.487*** | 0.397** | 0.318*** | 0.137** |
| | (0.125) | (0.030) | (0.092) | (0.185) | (0.029) | (0.067) |
| Baseline | 0.579 | 0.385 | 0.542 | 0.456 | 0.339 | 0.369 |
| Observations | 4,646 | 36,677 | 4,578 | 4,861 | 39,079 | 4,926 |
| R-squared | 0.039 | 0.014 | 0.029 | 0.009 | 0.009 | 0.026 |

*Notes:* The outcome variable is an indicator variable equal to one if the days of paid parental leave is at most half of the quota, i.e., 0 days before 1995, uptakes between 0 and 15 days for children born 1995- 2001, between 0 and 30 days for children born 2002-2015. Constraint is an indicator variable for having either physical (separated, sick or convicted) or financial (self-employment, unstable employment, high income share) constraints. Estimates from separate RD-estimations using a 6-month reform window on each side, triangular weights, and quadratic separate slopes. Robust standard errors in parentheses. *** p<0.01, **p<0.05, * p<0.1

Table 2 reports the regression estimates for having any constraint by income group for each reform. In both 1995 and 2002, the quotas reduced the probability of low uptake on average ($\gamma_2 < 0$), though the magnitude of the effect varies across income groups and reform years.[7] These coefficients capture the response of unconstrained fathers. For the 1995 reform, the decline in low uptake was largest among low-income fathers and smaller among those in the middle of the distribution, with no discernible change at the top. By contrast, the 2002 expansion produced little or no additional reduction in low uptake, consistent with the idea that most behavioral adjustments had already occurred following the first reform.

[7] Note that because the outcome is defined relative to the quota, a higher uptake following the reform does not necessarily render a negative average effect.

Across both reforms, constrained fathers responded similarly to unconstrained fathers. The interaction terms ($\gamma_1$) vary in sign across income groups and are statistically insignificant in all cases. This indicates that the quotas did not differentially alleviate the economic, physical, or normative barriers identified in the descriptive analysis. Supplementary analyses in the Appendix, Tables A1–A3, confirm that the pattern remains when constraints are disaggregated by type, effectively removing any concerns that the insignificant average effects are masking counteracting heterogeneity across constraints.

Overall, the 1995 and 2002 quotas increased fathers' average uptake but did not narrow the gap between constrained and unconstrained groups. Institutional reforms thus appear effective in shifting average behavior but less so in relaxing the underlying economic and normative constraints that sustain persistently low uptake.

## 5. The role of gender norms

Section 3 showed that economic considerations are consistently binding at the tails of the income distribution, while workplace norms have become an increasingly salient constraint for the majority of fathers over time. As noted in the previous section, earmarked leave did not narrow the gaps between constrained and unconstrained fathers. These findings suggest that stronger incentives alone are insufficient to shift behavior among the groups least likely to take leave, which motivates a closer examination of the normative factors sustaining persistently low uptake. Therefore, in this section we broaden our focus to social norm constraints operating beyond the workplace setting. First, we exploit regional variation in gender attitudes to assess how local cultural environments influence leave-taking. Second, we use quasi-experimental variation in sibling sex composition as a proxy for fathers' early-life exposure to gender-role environments.

### *5.1 Gender norms at regional level*

We begin by examining geographical variation in gender attitudes. Fathers' leave uptake differs markedly across Swedish municipalities (Appendix Figure A1), suggesting that local gender norms may shape the environment in which fathers make decisions. For example, regions with more egalitarian norms may foster environments where fathers feel greater social acceptance and encouragement in taking leave. In contrast, regions with more traditional norms may create social pressures or stigma that discourage fathers from doing so. Unlike workplace-specific norms analyzed earlier, regional variation captures a broader set of influences—household, community, and workplace—that jointly condition fathers' choices.

Following the literature that identifies the effects of gender norms on individual outcomes through cross-country variation (e.g., Guiso et al., 2008; Fogli & Fernandez, 2009; Hyde & Mertz, 2009; Aldén & Neuman, 2022), we extend this approach to exploit regional variation within Sweden. As proxies for local norms, we use county-level responses from the World Values Survey (WVS).

**Table 3. Regional gender norms and the probability of low parental leave uptake**

| | (1) *University is more important for a boy than for a girl* | (2) *Men make better political leaders than women do* | (3) *When jobs are scarce, men should have more right to a job than women* |
|---|---|---|---|
| | 1.261*** | 0.482*** | 1.476*** |
| | (0.282) | (0.115) | (0.287) |
| Observations | 255,921 | 255,921 | 255,921 |
| R-squared | 0.007 | 0.015 | 0.018 |

*Notes:* The table shows the relationship between local gender norms and the probability of taking at most half of the reserved parental leave days. Uptake is measured as the number of paid parental leave days during the child's first two years. Low uptake is defined as taking at most half of the reserved parental leave days. This corresponds to at most 15 days for children born before 2002 and 30 days for those born after. Gender norms are proxied by the share of county-level respondents agreeing with a given statement from the World Value Survey; higher values indicate more traditional beliefs. The 1996 WVS wave is used for children born before 2007; the 2006 WVS for children born 2007–2011, and the 2011 wave for children born after 2011. Columns 1–3 use three alternative statements as proxies. Controls include fathers' age at birth. Robust standard errors, clustered at birth municipality, in parentheses. *, **, and *** indicates statistical significance at the 10-, 5, and 1-percent level.

Our measure is the share of respondents who agree with statements reflecting traditional gender beliefs, with higher values indicating more traditional attitudes (see Web Appendix Table W4 for details).[8] Our central hypothesis is that traditional local norms are negatively associated with fathers' parental leave uptake. Table 3 presents the relationship between three local gender norm proxies and the likelihood of fathers taking at most half of the reserved parental leave days. Across all proxies, fathers in regions with more traditional gender norms are more likely to exhibit low uptake. A one standard deviation (std. = 0.039) increase in agreement with 'University is more important for a boy than for a girl' raises the probability of low uptake by 4.7 percentage points. The corresponding associations are 3.0 percentage points (std. = 0.063) for agreement with 'Men make better political leaders than women' and 5.2 percentage points (std. = 0.035) for 'When jobs are scarce, men should have more right to a job than women'.[9]

These descriptive results indicate that gender norms shape parental leave behavior. However, the associations are correlational and may reflect sorting, as fathers less inclined to take leave may settle in regions where leave-taking is less common or less socially supported. To obtain causal evidence on the role of gender norms, we exploit quasi-random variation in sibling-sex composition as a source of exogenous exposure to gendered environments.

*5.2 Gender norms at the individual level*

Parents of opposite-sex children tend to engage in more gender-differentiated parenting than those with same-sex children (e.g., McHale et al., 2003). Such parenting practices have been shown to shape individuals' gender conformity, particularly among women. Using Danish data,

[8] The WVS provides regional data only for three years (1996, 2006, and 2011) during our observation period, and only at the county level, see Inglehart et al (2014) for the dataset and www.worldvaluessurvey.org for codebooks. and detailed documentation. We include regions with at least 50 respondents per wave and restrict the sample to fathers with children born from 1997 onward to ensure that norms are measured before the child's birth. Due to the limited data, we abstract from analyzing trends. While relatively few respondents agree with the gender-norm statements, consistent with Sweden's relative gender equality, there is notable regional variation. The highest agreement is with the statement 'Men make better political leaders than women,' with an average agreement of 12 percent and a regional maximum of 29.2 percent. For the other two statements, average agreement is 4.5 and 5.1 percent, respectively, with significant regional variation (see Web Appendix Table W4 for details).

[9] This has been calculated, using estimates from column 1) and standard deviations from Web Appendix Table W5, as $\beta \times$ WVS_norm_std = 1.115 × 0.039 = 0.043.

Brenøe (2021) demonstrates that women with a younger brother, as opposed to a sister, are more likely to adopt traditional gender roles, reflected in occupational and partner choices.

Building on this evidence, we use the presence of a younger sister as a proxy for exposure to traditional gender norms and estimate its impact on fathers' parental leave uptake. We hypothesize that men who grew up with a sister are more likely to adhere to traditional gender norms than those who grew up with a brother.[10] Our identification strategy relies on the fact that the gender of the second-born child is effectively random, conditional on family size, allowing us to isolate the influence of gender norms from other confounding factors.

We restrict our analysis to fathers who are first-born sons in two-child families.[11] To ensure consistency in family structure, we further limit the sample to fathers who are the first-born child to both parents, with a sibling age gap of no more than four years, and exclude twin births.

To assess the effect of having a younger sister on father's low parental leave uptake, we estimate a linear probability model of the form:

$$y_i^{first-born} = \alpha_0 + \alpha_1 Sister_i^{second-born} + X_i'\beta + \varepsilon_i \qquad (4),$$

where $y_i^{first-born}$ equals one if the father takes at most half of the reserved parental leave days during the child's first two years and zero otherwise. $Sister_i^{second-born}$ is the key independent variable, equal to one if the second-born sibling is a sister and zero if the sibling is a brother. Thus, $\alpha_1$ captures the effect of having a younger sister, relative to having a younger brother, on the probability of having a low parental leave uptake. $X_i'$ is a vector of control variables including the father's birth year, the spacing to the second-born sibling, the father's county of

[10] Brenøe (2021) shows that parents of mixed-sex children engage in more gender-specific parenting than parents of same-sex children, with mothers spending relatively more time with daughters and fathers with sons. This pattern is in line with stronger transmission of gender-specific human capital and traditional gender norms in mixed-sex families.

[11] We focus on two-child families because sibling-sex compositions are more comparable when family size is held constant. As a robustness check, presented in Web Appendix W6, we also estimate regressions using a sample of families with at least two children, controlling for family size. Although this approach results in some loss of precision and somewhat smaller effect sizes, the estimates remain directionally consistent.

birth, the grandparents' age at the father's birth and their level of education. We use robust standard errors, clustered at the family level.[12]

To validate our proxy for exposure to gender norms, we first estimate first-stage regressions of sibling-sex composition on men's gender-typed educational and occupational choices, following Brenøe (2021). Full results and variable definitions are reported in Web Appendix W5 (Tables W5–W6). We find that fathers with a younger sister are more likely to enter male-dominated fields of study and work, consistent with sibling-sex composition shaping gender-differentiated upbringing. These effects are strongest among men from less-educated families, which aligns with evidence that more educated parents transmit more egalitarian norms (Geisler & Kreyenfeld, 2011). Because this heterogeneity may weaken the quasi-experimental variation, potentially generating effects that run in opposite directions, we present results stratified by grandparental education in the main analysis.

**Table 4: Effect of having an opposite-sex sibling on the probability of taking at most half of the reserved parental leave days**

| VARIABLES | (1) 1995-2015 | (2) 1995-2001 | (3) 2002-2009 | (4) 2010-2015 | (5) 1995-2015 | (6) 1995-2001 | (7) 2002-2009 | (8) 2010-2015 |
|---|---|---|---|---|---|---|---|---|
| Second-born sister | 0.002 | -0.003 | -0.001 | 0.011** | 0.001 | -0.001 | -0.003 | 0.012** |
| | (0.003) | (0.005) | (0.005) | (0.004) | (0.004) | (0.005) | (0.006) | (0.006) |
| Second-born sister x Higher | | | | | 0.001 | -0.005 | 0.004 | -0.002 |
| | | | | | (0.005) | (0.011) | (0.008) | (0.011) |
| Mean of dependent variable | 0.301 | 0.341 | 0.297 | 0.277 | 0.301 | 0.341 | 0.297 | 0.277 |
| Observations | 89,194 | 25,696 | 36,631 | 26,867 | 89,194 | 25,696 | 36,631 | 26,867 |
| R-squared | 0.022 | 0.022 | 0.027 | 0.040 | 0.019 | 0.019 | 0.024 | 0.035 |

*Notes:* The table shows the effect of having an opposite-sex sibling on fathers' probability of taking at most half of the reserved parental leave days during the child's first two years. Taking at most half of the reserved days corresponds to a maximum of 15 days for children born before 2002 and 30 days for those born in 2002 or later. All models control for the father's county and year of birth, spacing to the younger sibling, and grandparents' age at the father's birth and their level of education. In columns (1)–(4), grandparental education is included as indicators for level-by-field of education. In columns (5)–(8), grandparental education is defined as an indicator equal to one if at least one grandparent has a university degree and no grandparent has only primary education ("Higher"). Robust standard errors, clustered at the birth county level, are reported in parentheses. *, **, and *** indicate statistical significance at the 10-, 5-, and 1-percent levels, respectively.

[12] Appendix Table A4 presents descriptive statistics for pre-determined childhood environment characteristics of fathers. As expected under the assumption of random sibling sex, there is little systematic variation in these characteristics between fathers with a younger brother and those with a younger sister. Although some differences in grandparental age and education are statistically significant, they are small in magnitude. These variables are therefore included as controls in our regressions.

Table 4 presents estimates of the effect of having a younger sister on fathers' low uptake. For the full period 1995–2015, sibling-sex composition has no significant effect on the likelihood of low uptake (column 1). The pooled estimate may mask important changes over time. Our reform analysis shows that quotas increased uptake uniformly but did not narrow constraint-related gaps, and our descriptive evidence indicates that workplace norms have become increasingly influential across cohorts. These patterns suggest that the role of gender norms may have strengthened over time, motivating an examination of cohort-specific effects.

Indeed, we find notable heterogeneity across birth cohorts.[13] For fathers of children born before 2010, the estimates are close to zero and statistically insignificant (see columns 2–3). By contrast, among first-time fathers in the most recent cohort (2010–2015), having a younger sister increases the probability of low uptake: these fathers are 1.1 percentage points (3.7 percent) more likely to take no more than half of their reserved leave days (see column 4).

Columns 5–8 further investigate whether the effect of sibling-sex composition varies by grandparental education. Across all cohorts, the interaction terms are small and statistically insignificant, indicating no meaningful difference in the effect of sibling-sex composition between men from higher- and lower-educated families. Importantly, the positive effect observed for the 2010–2015 cohort remains virtually unchanged once the interaction is included.[14]

To reinforce these findings, we also assess whether the first-stage relationships between sibling-sex composition and gender-typed educational and occupational choices vary across cohort (see Web Appendix Table W6). These relationships remain stable over time, indicating that the sibling-sex proxy consistently captures exposure to gender norms. Combined with the

[13] Due to limited statistical power at the single-year level, we pool birth years into cohorts.

[14] We conduct several robustness checks reported in the Web Appendix Table W6. These include relaxing the two-child family restriction, controlling for family size, and adjusting for the child's gender. Across these specifications, the results remain broadly consistent, with the positive effect of having a younger sister persisting for the 2010–2015 cohort, though estimates remain small and statistically insignificant for earlier cohorts.

declining share of fathers with low uptake, this pattern suggests that the remaining low-uptake group has become increasingly selected into more traditional gender norms. In other words, the shrinking group of fathers who take little leave appears to be more norm-conservative than earlier cohorts, which helps explain why the influence of gender norms strengthens in the most recent birth cohort.

Our earlier analyses indicate that gender norms matter most for fathers in the middle of the income distribution (deciles 2–9), while economic constraints dominate at the top and bottom. Restricting the analysis to fathers in the middle of the income distribution (deciles 2–9), where gender norms are most relevant, slightly strengthens the effect for the most recent cohort, while estimates for earlier cohorts remain small and insignificant (Appendix Table A5). This pattern supports the idea that among fathers for whom economic and career constraints are less binding, parental leave decisions are more strongly driven by gender norms, and that this has become increasingly the case in recent cohorts.

## 6. Conclusion

Despite extensive reforms aimed at increasing fathers' involvement in early childcare, many fathers across advanced economies still make limited use of paid parental leave. We examine a setting where financial and eligibility barriers are low: employed, Swedish-born first-time fathers living with their child who receive generous earnings-related benefits and for whom unused reserved days are lost by the household. Focusing on these fathers provides a sharp test of the factors that constrain uptake when neither access nor affordability is binding.

Three robust findings emerge. First, low uptake follows a persistent U-shaped income gradient, being most common among fathers at the bottom and top of the earnings distribution, and least common among those in the middle. Second, the drivers of low uptake differ systematically across the income distribution: economic constraints such as self-employment and unstable work bind at the bottom, high opportunity costs and breadwinner roles matter at the top, while workplace and cultural norms are particularly salient among the middle-income fathers—who constitute the majority of fathers. Third, these constraints have become increasingly important over time, but in different ways across the income distribution: workplace norms have grown in relevance for middle-income fathers, whereas opportunity costs have intensified at the top.

Quota reforms in 1995 and 2002 raised fathers' parental leave uptake but did not narrow constraint-related gaps in low uptake. Constrained and unconstrained fathers responded similarly to both reforms, suggesting that expanding earmarked leave alone is insufficient to relax the underlying economic and normative barriers that sustain persistently low uptake. Our quasi-experimental analysis using sibling-sex composition indicates that gender-role norms influence fathers' parental leave decisions. Fathers exposed to more traditional gender-role environments in childhood are more likely to abstain from leave, with effects concentrated among recent cohorts. This pattern suggests that normative constraints are persistent and have become increasingly influential, even among fathers for whom financial and economic barriers are minimal.

Finally, the analysis excludes fathers with weak labor market attachment, who remain underrepresented in leave uptake and largely untouched by quota reforms. Their exclusion underscores that universal entitlements do not ensure universal use, and that socioeconomic disadvantage continues to shape access to policy benefits.

Taken together, our findings suggest that while parental leave reforms have expanded fathers' uptake, they have not altered the underlying factors that sustain unequal caregiving. Further progress requires more than improved access and affordability. Reducing economic barriers remains necessary, but durable change will also depend on addressing normative constraints within households and workplaces that continue to discourage fathers from taking leave. Lasting change may require interventions that target workplace practices and gender norms, not only benefit design. Future work should explore how such policies interact, complementing financial incentives with measures that shift expectations within families and firms.

**Data availability.**
All analyses are based on confidential Swedish administrative register data accessed through Statistics Sweden and the Swedish Social Insurance Agency. The data can be obtained by filing a request with Statistics Sweden, see https://www.scb.se/en/services/guidance-for-researchers-and-universities/. To obtain the data used in the paper one must also apply for permission from the Swedish Ethical Review Authority at https://etikprovningsmyndigheten.se/. All data processing is conducted on secure servers at Statistics Sweden via remote terminal access. We are happy to provide the full replication code to qualified researchers upon request.

**Ethical statement.**
The study relies exclusively on anonymized administrative register data. No human subjects were contacted, and informed consent is not required under Swedish law for this type of research. This project was approved by the Swedish Research Ethics Board (Event No 2018/108-31/5), and the project complies with applicable data protection and ethical regulations.

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## APPENDIX FIGURES AND TABLES

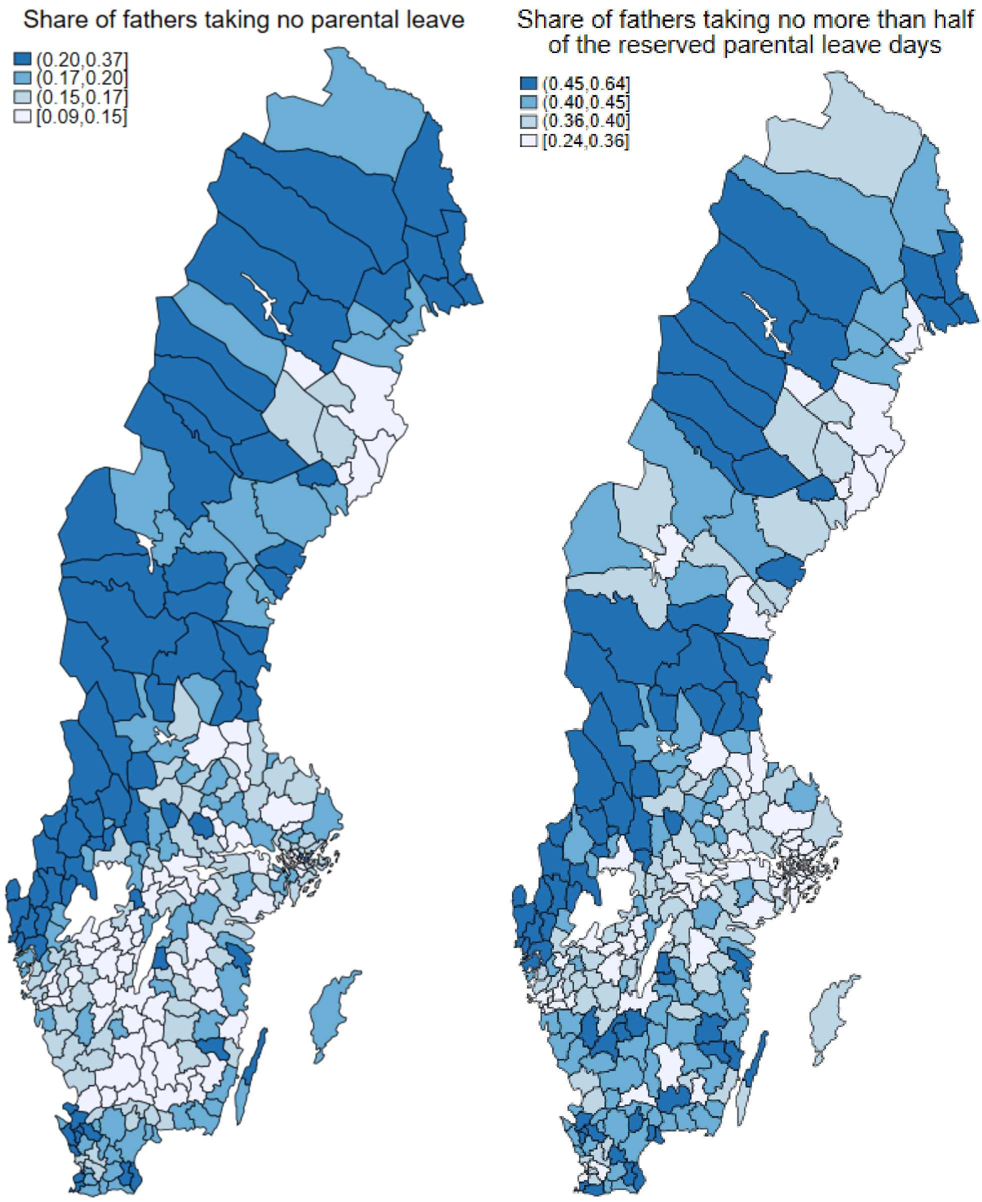


**Figure A1. Geographical variation in share of fathers taking no or at most half of the reserved parental leave days during the child's first two years old, 1995–2015.**

**Table A1. Reform analysis on parental leave uptake using regression discontinuity, economic constraints.**

| | (1) | (2) | (3) | (4) | (5) | (6) |
|---|---|---|---|---|---|---|
| | | Reform 1995 | | | Reform 2002 | |
| | Decile 1 | Decile 2-9 | Decile 10 | Decile 1 | Decile 2-9 | Decile 10 |
| TreatedXreformXconstr. | 0.030 | 0.059 | -0.164 | 0.181 | -0.077 | -0.113 |
| | (0.224) | (0.067) | (0.192) | (0.232) | (0.065) | (0.185) |
| TreatedXreform | -0.225 | -0.102** | -0.058 | -0.254 | 0.093** | 0.112 |
| | (0.192) | (0.041) | (0.134) | (0.205) | (0.042) | (0.138) |
| TreatedXconstr. | 0.113 | 0.054 | 0.207 | -0.176 | 0.025 | 0.203* |
| | (0.147) | (0.049) | (0.135) | (0.173) | (0.046) | (0.122) |
| ReformXconstr. | 0.018 | -0.056 | 0.078 | -0.128 | 0.010 | -0.162 |
| | (0.156) | (0.051) | (0.147) | (0.176) | (0.047) | (0.139) |
| Constrained | -0.004 | 0.081** | -0.033 | 0.061 | -0.002 | 0.143 |
| | (0.104) | (0.036) | (0.103) | (0.132) | (0.034) | (0.095) |
| Treated | 0.041 | -0.067** | -0.286*** | 0.273* | -0.037 | -0.078 |
| | (0.122) | (0.029) | (0.091) | (0.154) | (0.030) | (0.083) |
| Reform | -0.014 | 0.007 | 0.054 | 0.147 | -0.043 | 0.206* |
| | (0.130) | (0.032) | (0.107) | (0.157) | (0.031) | (0.107) |
| Constant | 0.573*** | 0.353*** | 0.527*** | 0.340*** | 0.343*** | 0.188*** |
| | (0.084) | (0.022) | (0.072) | (0.120) | (0.022) | (0.069) |
| | | | | | | |
| Observations | 4,646 | 36,677 | 4,578 | 4,861 | 39,079 | 4,926 |
| R-squared | 0.039 | 0.015 | 0.032 | 0.006 | 0.004 | 0.026 |

*Notes:* The outcome variable is an indicator variable taking 1 if the days of parental leave is at most half of the quota. i.e., 0 days before 1995, uptakes between 0 and 15 days for children born 1995- 2001, between 0 and 30 days for children born 2002-2015. Constraint is an indicator variable for having either physical (separated, sick or convicted) or financial (self-employment, unstable employment, high income share) constraints. All estimations are conditional on the parents living together either the year of birth or the year after. Estimates from separate RD-estimations using a 6-month reform window on each side, triangular weights, and quadratic separate slopes. Robust standard errors in parenthesis. *** p<0.01, **p<0.05, * p<0.1

**Table A2. Reform analysis on parental leave uptake using regression discontinuity, physical constraints.**

| | (1) | (2) | (3) | (4) | (5) | (6) |
|---|---|---|---|---|---|---|
| | | Reform 1995 | | | Reform 2002 | |
| | Decile 1 | Decile 2-9 | Decile 10 | Decile 1 | Decile 2-9 | Decile 10 |
| TreatedXreformXconstr. | 0.573* | -0.178 | -0.147 | 0.295 | 0.048 | -0.332 |
| | (0.321) | (0.136) | (0.553) | (0.288) | (0.138) | (0.517) |
| TreatedXreform | -0.251** | -0.055 | -0.124 | -0.142 | 0.055* | 0.050 |
| | (0.104) | (0.034) | (0.098) | (0.102) | (0.033) | (0.096) |
| TreatedXconstr. | -0.547*** | 0.030 | 0.159 | -0.371* | -0.033 | -0.139 |
| | (0.196) | (0.095) | (0.334) | (0.212) | (0.100) | (0.385) |
| ReformXconstr. | -0.247 | 0.113 | 0.396 | -0.444** | 0.009 | 0.364 |
| | (0.229) | (0.106) | (0.342) | (0.212) | (0.105) | (0.397) |
| Constrained | 0.358*** | 0.023 | -0.204 | 0.330** | 0.089 | 0.279 |
| | (0.121) | (0.074) | (0.265) | (0.151) | (0.076) | (0.297) |
| Treated | 0.175** | -0.051** | -0.193*** | 0.180** | -0.024 | 0.045 |
| | (0.071) | (0.024) | (0.069) | (0.073) | (0.024) | (0.067) |
| Reform | 0.023 | -0.022 | 0.083 | 0.096 | -0.039 | 0.106 |
| | (0.075) | (0.026) | (0.075) | (0.074) | (0.024) | (0.070) |
| Constant | 0.538*** | 0.384*** | 0.516*** | 0.351*** | 0.337*** | 0.267*** |
| | (0.053) | (0.018) | (0.052) | (0.054) | (0.017) | (0.050) |
| Observations | 4,646 | 36,677 | 4,578 | 4,861 | 39,079 | 4,926 |
| R-squared | 0.040 | 0.009 | 0.026 | 0.008 | 0.004 | 0.015 |

*Notes:* The outcome variable is an indicator variable equal to one if the days of parental leave is at most half of the quota. i.e., 0 days before 1995, uptakes between 0 and 15 days for children born 1995- 2001, between 0 and 30 days for children born 2002-2015. Constraint is an indicator variable for having either physical (separated, sick or convicted) or financial (self-employment, unstable employment, high income share) constraints. All estimations are conditional on the parents living together either the year of birth or the year after. Estimates from separate RD-estimations using a 6-month reform window on each side, triangular weights, and quadratic separate slopes. Robust standard errors in parenthesis. *** p<0.01, **p<0.05, * p<0.1

**Table A3. Reform analysis on parental leave uptake using regression discontinuity, social constraints.**

| | (1) | (2) | (3) | (4) | (5) | (6) |
|---|---|---|---|---|---|---|
| | | Reform 1995 | | | Reform 2002 | |
| | Decile 1 | Decile 2-9 | Decile 10 | Decile 1 | Decile 2-9 | Decile 10 |
| TreatedXreformXconstr. | 0.191 | -0.043 | -0.050 | -0.046 | 0.027 | 0.189 |
| | (0.196) | (0.066) | (0.204) | (0.191) | (0.066) | (0.226) |
| TreatedXreform | -0.267* | -0.048 | -0.124 | -0.094 | 0.047 | 0.016 |
| | (0.148) | (0.045) | (0.117) | (0.141) | (0.040) | (0.107) |
| TreatedXconstr. | -0.359*** | 0.006 | -0.081 | 0.110 | 0.016 | -0.074 |
| | (0.132) | (0.047) | (0.143) | (0.138) | (0.047) | (0.154) |
| ReformXconstr. | -0.179 | 0.009 | -0.062 | -0.040 | -0.036 | -0.058 |
| | (0.140) | (0.050) | (0.157) | (0.141) | (0.048) | (0.161) |
| Constrained | 0.332*** | 0.042 | 0.140 | 0.068 | 0.090*** | 0.067 |
| | (0.096) | (0.035) | (0.108) | (0.102) | (0.035) | (0.113) |
| Treated | 0.295*** | -0.050 | -0.159* | 0.082 | -0.031 | 0.061 |
| | (0.099) | (0.032) | (0.082) | (0.103) | (0.029) | (0.075) |
| Reform | 0.067 | -0.019 | 0.121 | 0.068 | -0.019 | 0.125 |
| | (0.106) | (0.034) | (0.089) | (0.105) | (0.030) | (0.079) |
| Constant | 0.413*** | 0.366*** | 0.461*** | 0.351*** | 0.302*** | 0.255*** |
| | (0.070) | (0.024) | (0.063) | (0.077) | (0.022) | (0.057) |
| Observations | 4,646 | 36,677 | 4,578 | 4,861 | 39,079 | 4,926 |
| R-squared | 0.046 | 0.012 | 0.028 | 0.017 | 0.015 | 0.017 |

*Notes:* The outcome variable is an indicator variable taking 1 if the days of parental leave is at most half of the quota. i.e., 0 days before 1995, uptakes between 0 and 15 days for children born 1995- 2001, between 0 and 30 days for children born 2002-2015. Constraint is an indicator variable for having either physical (separated, sick or convicted) or financial (self-employment, unstable employment, high income share) constraints. All estimations are conditional on the parents living together either the year of birth or the year after. Estimates from separate RD-estimations using a 6-month reform window on each side, triangular weights, and quadratic separate slopes. Robust standard errors in parenthesis. *** p<0.01, **p<0.05, * p<0.1

**Table A4. Descriptive statistics on parental leave uptake and childhood environment of first-born fathers**

| | (1) | (2) | (3) |
|---|---|---|---|
| | Second-born | | |
| | Sister | Brother | t-test (p-value) |
| *Parental leave outcomes* | | | |
| At most half of reserved PL days | 0.305 | 0.302 | 0.352 |
| | (0.461) | (0.459) | |
| *Occupational and educational outcomes* | | | |
| Male share in occupation | 0.711 | 0.707 | 0.023 |
| | (0.213) | (0.213) | |
| STEM occupation | 0.258 | 0.260 | 0.427 |
| | (0.386) | (0.389) | |
| Male-dominated occupation | 0.429 | 0.422 | 0.016 |
| | (0.423) | (0.421) | |
| Male share in education | 0.674 | 0.669 | 0.017 |
| | (0.269) | (0.270) | |
| STEM education | 0.486 | 0.479 | 0.172 |
| | (0.500) | (0.500) | |
| Male-dominated education | 0.484 | 0.478 | 0.038 |
| | (0.500) | (0.500) | |
| *Pre-determined characteristics* | | | |
| Spacing between siblings (years) | 2.9 | 2.9 | 0.001 |
| | (0.9) | (0.9) | |
| Grandmother's age at father's birth | 24.9 | 25.1 | 0.000 |
| | (3.8) | (3.9) | |
| Grandfather's age at father's birth | 27.4 | 27.5 | 0.000 |
| | (4.3) | (4.4) | |
| Grandmother's years of education | 11.4 | 11.4 | 0.425 |
| | (2.3) | (2.3) | |
| Grandfather's years of education | 11.4 | 11.4 | 0.081 |
| | (2.5) | (2.5) | |
| Observations | 45,675 | 43,519 | |

*Notes:* The table reports descriptive statistics for first-born fathers in two-child families, shown separately for those with a second-born sister (Column 1) and those with a second-born brother (Column 2). Column 3 reports p-values from t-tests of differences between the two groups. Parental leave uptake refers to the number of paid parental leave days taken during the child's first two years. Taking at most half of the reserved days corresponds to a maximum of 15 days for children born before 2002 and 30 days for those born in 2002 or later. Standard deviations are reported in parentheses.

**Table A5: Effect of sibling sex on probability to take at most half of the reserved parental leave days– income decile 2–9**

| VARIABLES | (1)<br>1995-<br>2015 | (2)<br>1995-<br>2001 | (3)<br>2002-<br>2009 | (4)<br>2010-<br>2015 | (5)<br>1995-<br>2015 | (6)<br>1995-<br>2001 | (7)<br>2002-<br>2009 | (8)<br>2010-<br>2015 |
|---|---|---|---|---|---|---|---|---|
| Second-born sister | 0.003 | -0.005 | 0.002 | 0.014*** | 0.004 | -0.002 | 0.001 | 0.016** |
| | (0.003) | (0.005) | (0.006) | (0.004) | (0.003) | (0.006) | (0.006) | (0.006) |
| Second-born sister x Higher grandparental education | | | | | -0.001 | -0.009 | 0.002 | -0.007 |
| | | | | | (0.005) | (0.013) | (0.008) | (0.011) |
| Observations | 79,215 | 23,037 | 31,964 | 24,214 | 79,215 | 23,037 | 31,964 | 24,214 |
| R-squared | 0.022 | 0.024 | 0.028 | 0.040 | 0.019 | 0.020 | 0.024 | 0.035 |

*Notes:* The table shows the effect of having an opposite-sex sibling on fathers' probability of taking at most half of the reserved parental leave days during the child's first two years. Taking at most half of the reserved days corresponds to a maximum of 15 days for children born before 2002 and 30 days for those born in 2002 or later. The sample includes fathers in income deciles 2–9. All models control for the father's county and year of birth, spacing to the younger sibling, and grandparents' age at the father's birth and their level of education. In columns (1)–(4), grandparental education is included as indicators for level-by-field of education. In columns (5)–(8), grandparental education is defined as an indicator equal to one if at least one grandparent has a university degree and no grandparent has only primary education ("Higher"). Robust standard errors, clustered at the birth county level, are reported in parentheses. *, **, and *** indicate statistical significance at the 10-, 5-, and 1-percent levels, respectively.

# WEB APPENDIX

## Father but not caregivers

**Lina Aldén, Anne Boschini and Malin Tallås Ahlzén**

## Web Appendix W1. The Swedish parental insurance and trends in fathers' paid parental leave uptake

In Sweden, both parents have been entitled to parental leave benefits since 1974. For many years, the shared leave was used almost exclusively by mothers, but fathers' uptake has increased markedly in recent decades (see Figure W1, left panel). To promote a more gender-equal division, several reforms have introduced non-transferable quotas—leave days reserved for each parent. The first quota was implemented in 1995, earmarking one month of paid leave for each parent while keeping the total number of leave days unchanged. A second month was added in 2002.[15] The 2002 reform also expanded the number of shared leave days, which left maternal uptake largely unaffected and may have diluted the reform's effect on fathers' behavior (Swedish Social Insurance Agency, 2014). As a result, the 2002 quota is expected to have had a weaker impact on fathers' uptake than the reform of 1995.

In 2012, parents were given the option to take up to 30 days of leave simultaneously during the child's first year—so-called "double days." This reform encouraged more fathers to take at least some leave early in the child's life. However, evaluations suggest that these days are partially substituted for reserved quota days, which cannot be taken as double days (Fahlén & Bjurström, 2018).

Since the introduction of reserved parental leave days for each parent, fathers' total leave uptake has increased substantially—from an average of 47 days for children born in 1994 to 125 days for those born in 2010. However, this growth is less pronounced when focusing on the child's early years. During the first year of life, fathers took approximately 22 days of leave in 1994, compared to about one month in 2016. A substantial portion of leave continues to be taken after the child's second birthday, which is partly explained by the generous flexibility in the system: until 2014, parental leave could be used until the child turned eight years old, and

[15] A third month of reserved parental leave was introduced in 2016, which falls outside the study period.

between 2014 and 2016, up to the day the child started school. Moreover, increases in average uptake mask considerable heterogeneity. While some fathers have expanded their use of leave over time, a non-negligible share continues to take little or no paid leave. This variation is illustrated in the right-hand panel of Figure W1, which shows trends in the share of first-time fathers taking no leave, less than half of the reserved quota, or at most the full quota during the child's first two years. Overall, the figure highlights a striking stability in the share of fathers taking no or very limited leave—despite repeated policy efforts to increase fathers' involvement in early childcare.

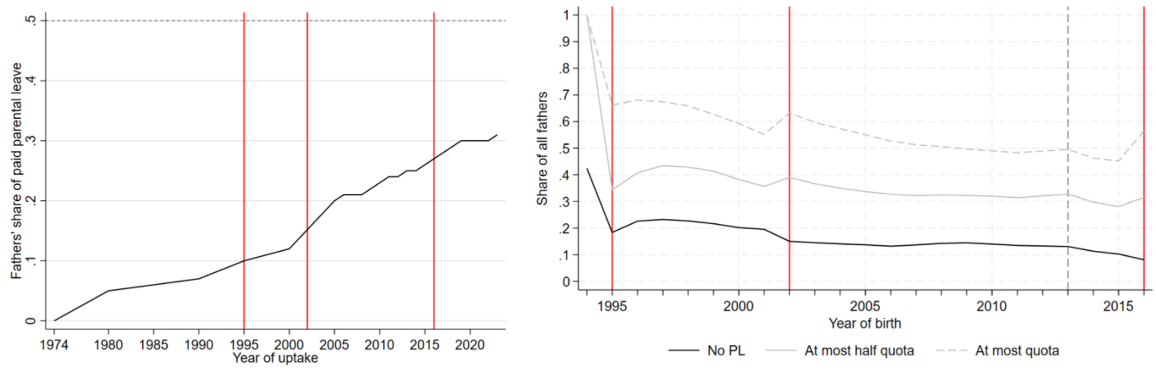


**Figure W1. The share of fathers by degree of parental leave uptake for children by year of birth**
Note: Panel A shows the average time trend in fathers' uptake of paid parental leave as a share of the family's total parental leave for a child. Panel B reports the shares of fathers who do not take any leave, who take at most half of the reserved amount of leave (15 days from 1995–2001 and 30 days from 2002), and who take at most the reserved amount of leave (30 days from 1995–2001 and 60 days from 2002). Red vertical lines indicate the introduction of the quotas, and the gray dashed line indicates the introduction of 'double-days' in parental leave insurance. Due to data availability, we include only births until July 2016.

Swedish parents are currently entitled to 18 months of job protection and 480 days of paid parental leave benefits. Of these, 390 days are income-based, replacing approximately 80 percent of previous earnings up to a cap, while the remaining 90 days are paid at a low flat rate. Parents without sufficient earnings to qualify for the income-based benefit receive a basic flat-rate allowance. Prior to 2002, only 360 days were income-based, and both the earnings ceiling and flat-rate amounts have been adjusted upward several times over the study period. In 2016, for example, the income cap was set at 10 price base amounts (PBA), the flat-rate benefit at SEK 180 per day, and the basic benefit at SEK 250 per day (Swedish Social Insurance Agency, 2022). In addition to the 480 parental leave days, fathers are entitled to 10 days of birth-related paternity leave, to be taken within 60 days of the child's birth. These 10 days are not included in our measures of parental leave uptake.

Although the structure of the benefit system has evolved, compensation levels have remained comparatively generous. When parental leave was introduced, the replacement rate was 90 percent. This was temporarily lowered during the economic crisis of the 1990s and currently stands at 77.6 percent (Swedish Social Insurance Agency, 2024). For parents without earnings, the flat-rate benefit provides not only financial support but also an incentive to establish labor market attachment before childbirth. In addition to state benefits, many employed parents—especially those covered by collective agreements—receive supplementary "parental pay" from their employers, which further mitigates income loss during leave.

Still, fathers' uptake of their reserved parental leave varies substantially across the earnings distribution. Figure W2 displays the share of fathers who used at most half of their reserved quota during the child's first two years, by pre-birth earnings decile, for two periods: 1995–2000 and 2010–2015. Although the overall level of limited uptake has fallen over time, the figure reveals clear and persistent differences across income groups. In the more recent period, the likelihood of taking only part of the reserved leave declines steadily with earnings, indicating that higher-income fathers have increased their uptake more than those with lower earnings. This pattern highlights the importance of considering heterogeneity in constraints and incentives across the income distribution when evaluating policies designed to promote fathers' engagement in early childcare.

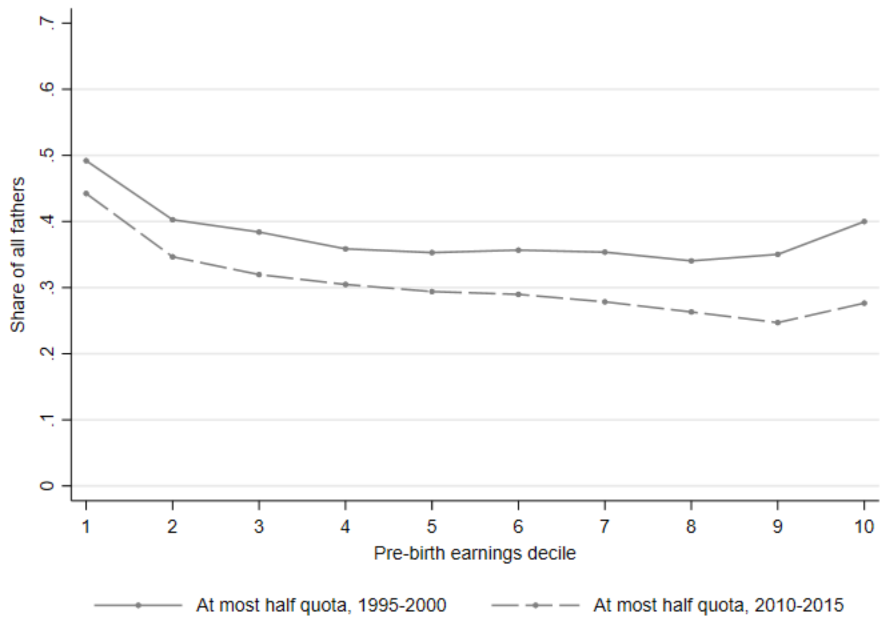


**Figure W2. Share of fathers taking at most half of their reserved parental leave, by pre-birth earnings decile: children born 1995–2000 and 2010–2015**

To assess how different factors contribute to low parental leave uptake across the earnings distribution, we categorize fathers into three groups based on their pre-birth earnings: decile 1 (lowest earners), deciles 2 through 9 (middle earners), and decile 10 (highest earners)—see Figure W3. The rationale for this grouping is both empirical and interpretive. As shown in the figure, the probability of taking at most half of the reserved quota is relatively homogeneous among fathers in deciles 2 to 9, suggesting that these fathers face similar constraints and behaviors with respect to leave-taking. In contrast, the patterns observed in the lowest and highest deciles are distinct: low-income fathers consistently exhibit higher non-uptake, while uptake among top earners diverges in both level and trend over time.

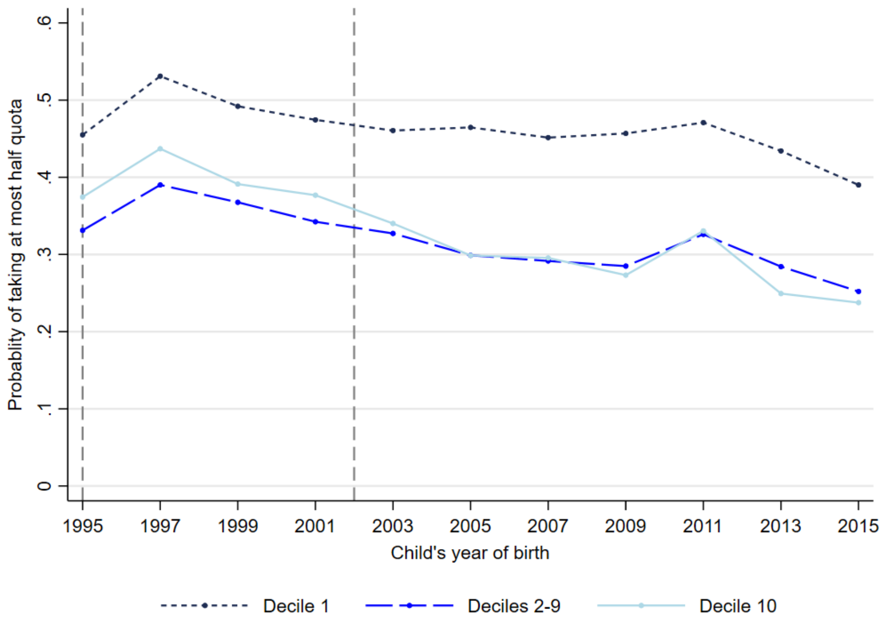


**Figure W3. The share of fathers by degree of parental leave uptake and earnings group for children born 1994–2016**
*Note:* The figure shows the share of fathers who take at most half of the quota days of paid parental leave, divided into three earnings groups based on their prebirth earnings rank.

## Web Appendix W2. Descriptive statistics on the analytical sample of fathers and their constraints

The analysis in this paper focuses on a selected sample of fathers, drawn from the full population of fathers with children born in Sweden between 1995 and 2015 for whom at least one parent took parental leave. To focus on fathers with low eligibility barriers for leave-taking, minimize confounding factors related to labor market attachment and family structure, and ensure complete data, we apply several sample restrictions: we exclude foreign-born fathers, students, individuals without a registered workplace, fathers of higher-parity births, and those not cohabiting with the mother at the time of childbirth. Figure W4 shows the share of fathers included in the analytical sample by pre-birth income decile of the full population of fathers of children born in 1995–2015. Sampling rates are very low in the bottom decile—reflecting the exclusion of fathers with weak or no labor market attachment—but remain relatively stable across the middle of the distribution, before declining slightly again in the top decile.

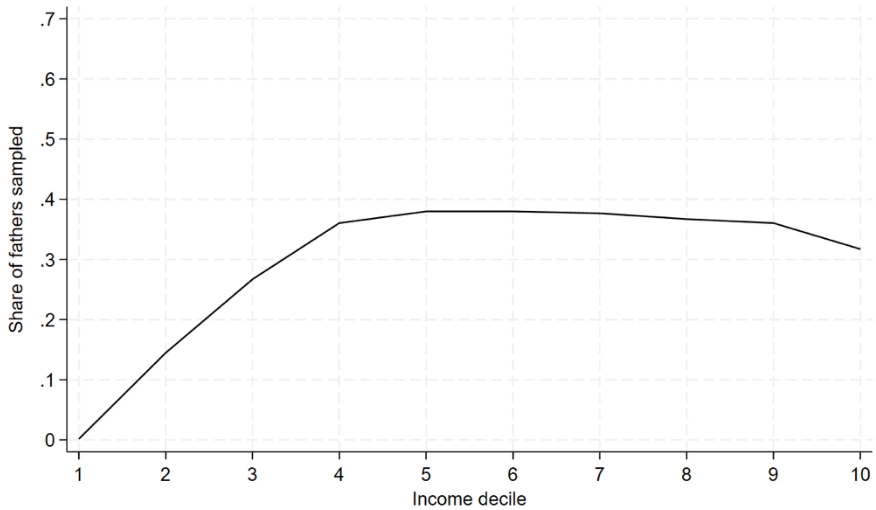


**Figure W4. Share of fathers included in the analytic sample, by pre-birth earnings decile of the full population of fathers**
*Note:* The figure shows the share of fathers included in the analytical sample by the pre-birth earnings distribution of all fathers of children born in 1995–2015. Pre-birth earnings are measured in the calendar year before childbirth and include annual labor income from wage- and self-employment as well as capital income.

Figure W5 instead compares the prevalence of low parental leave uptake—defined as using at most half of the reserved quota—between the full population and the analytic sample. As expected, the share of low-uptake fathers is somewhat lower in the analytic sample, particularly in the lower deciles, due to the exclusion of groups more likely to forgo leave entirely (e.g., students or non-resident fathers). However, the overall pattern across the income distribution is preserved, supporting the validity of the sample for analyzing variation in leave-taking behavior.

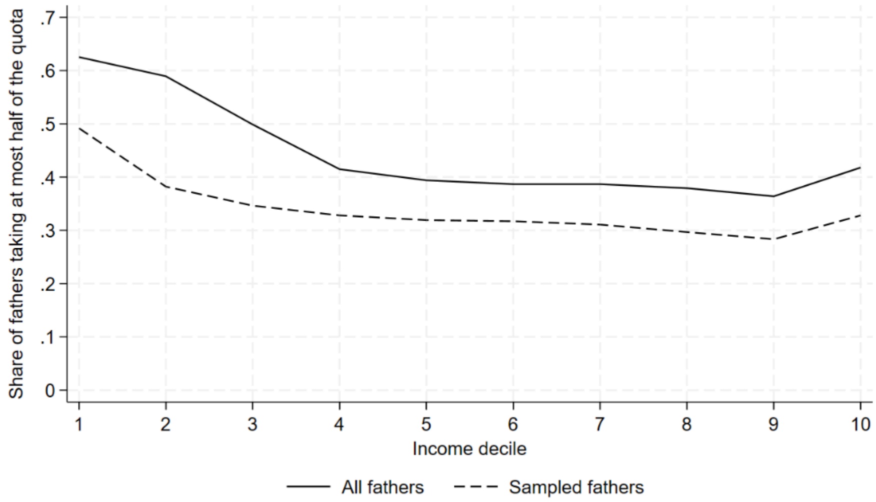


**Figure W5. Share of fathers taking at most half of the reserved parental leave quota: full population vs. analytic sample**
*Note:* The figure shows the share of fathers of children born in 1995–2015 who take at most half of the reserved parental leave quota, both in the analytical sample and in the full population of fathers, by fathers' pre-birth earnings distribution. Pre-birth earnings are measured in the calendar year before childbirth and include annual labor income from wage- and self-employment as well as capital income. The parental leave uptake refers to the number of paid parental leave days during the child's first two years. Taking at most half of the reserved quota corresponds to 0–15 days for fathers of children born in 1995–2001 and to 0–30 days for fathers of children born in 2002–2015.

Table W1 presents descriptive statistics for all fathers in our analytical sample and for subgroups by parental leave uptake and pre-birth earnings decile (based on fathers in the analytical sample). Columns (1) and (2) report averages for all fathers and those taking at most half of the reserved parental leave quota. Columns (3)–(5) display statistics by pre-birth earnings decile for the full analytical sample, while columns (6)–(8) show the same breakdown for low-uptake fathers.

Most fathers were aged 25–34 at childbirth, with the largest share between 25 and 29. Fathers taking at most half of the quota are slightly overrepresented in the youngest (under 25) and oldest (40+) age groups. Educational attainment differs substantially across groups: 30 percent of all fathers held a university degree, compared to 19 percent among low-uptake fathers. The share with only primary education is notably higher among low-uptake fathers (10 percent) than among all fathers in the sample (8 percent), and these patterns vary considerably across the income distribution. In the bottom decile, nearly 19 percent of low-uptake fathers have only a primary education, while over 50 percent of those in the top decile hold a university degree.

**Table W1. Descriptive statistics of fathers in the analytical sample by parental leave uptake and pre-birth earnings decile**

| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) |
|---|---|---|---|---|---|---|---|---|
| | | | | **All** | | **At most half the quota** | | |
| | **All** | **At most half quota** | Decile 1 | Deciles 2-9 | Decile 10 | Decile 1 | Deciles 2-9 | Decile 10 |
| *Age at birth* | | | | | | | | |
| <20 | 0.007 | 0.012 | 0.043 | 0.003 | 0.000 | 0.050 | 0.005 | 0.000 |
| 20-24 | 0.153 | 0.195 | 0.318 | 0.151 | 0.009 | 0.324 | 0.195 | 0.014 |
| 25-29 | 0.392 | 0.387 | 0.335 | 0.424 | 0.199 | 0.322 | 0.421 | 0.219 |
| 30-34 | 0.310 | 0.274 | 0.204 | 0.301 | 0.482 | 0.198 | 0.265 | 0.457 |
| 35-39 | 0.107 | 0.101 | 0.074 | 0.10 | 0.234 | 0.077 | 0.088 | 0.227 |
| 40- | 0.031 | 0.032 | 0.026 | 0.026 | 0.076 | 0.029 | 0.026 | 0.082 |
| *Education* | | | | | | | | |
| Primary | 0.077 | 0.101 | 0.170 | 0.073 | 0.020 | 0.186 | 0.094 | 0.031 |
| Secondary | 0.619 | 0.707 | 0.691 | 0.642 | 0.357 | 0.712 | 0.737 | 0.467 |
| University | 0.303 | 0.190 | 0.135 | 0.284 | 0.624 | 0.097 | 0.167 | 0.501 |
| Missing info | 0.001 | 0.002 | 0.004 | 0.001 | 0.000 | 0.005 | 0.001 | 0.001 |
| *Child characteristics* | | | | | | | | |
| Child is a boy | 0.515 | 0.513 | 0.512 | 0.516 | 0.514 | 0.512 | 0.513 | 0.512 |
| Observations | 530,875 | 177,270 | 53,098 | 424,702 | 53,075 | 24,623 | 135,249 | 17,398 |

*Notes:* The table reports descriptive statistics for our analytical sample of fathers of children born in 1995–2015, by parental leave uptake and pre-birth earnings decile. Parental leave uptake refers to the number of paid parental leave days during the child's first two years. Taking at most half of the reserved quota corresponds to 0–15 days for fathers of children born in 1995–2001 and 0–30 days for those of children born in 2002–2015. Pre-birth earnings are measured in the calendar year before childbirth and include annual labor income from wage- and self-employment as well as capital income. Educational attainment is measured in the year before childbirth; primary education corresponds to at most nine years of schooling.

Table W2 reports the share of fathers facing economic, physical, and norm-related constraints among those who took at most half of their reserved parental leave quota. The results are shown separately by pre-birth earnings decile and birth cohort period (1995–2000 and 2010–2015).

Definitions of the constraints are provided in Section 2.3 of the paper. The table shows that constraints are unequally distributed across the income distribution, with low-income fathers (decile 1) consistently being more constrained. Among this group, over 90 percent face at least one constraint in both periods. Although some individual indicators—such as unemployment and separation—have declined slightly over time, the overall prevalence of economic and physical constraints remains remarkably stable. Normative constraints, proxied by exposure to traditional workplace environments, follow a similarly persistent socioeconomic gradient.

**Table W2. Prevalence of constraints among low-uptake fathers, by earnings decile and period**

| | 1995-2000 | | | 2010-2015 | | |
|---|---|---|---|---|---|---|
| | Decile 1 | Deciles 2-9 | Decile 10 | Decile 1 | Deciles 2-9 | Decile 10 |
| **Economic constraints** | | | | | | |
| Changing job | 0.160 | 0.138 | 0.252 | 0.272 | 0.169 | 0.190 |
| Previous unemployed | 0.434 | 0.233 | 0.027 | 0.259 | 0.095 | 0.021 |
| Self-employed | 0.261 | 0.045 | 0.053 | 0.262 | 0.086 | 0.146 |
| High share of household income | 0.132 | 0.118 | 0.423 | 0.167 | 0.176 | 0.463 |
| Share with any economic constraints | 0.809 | 0.476 | 0.610 | 0.776 | 0.453 | 0.637 |
| Average number of economic constraint | 0.987 | 0.534 | 0.756 | 0.960 | 0.527 | 0.819 |
| **Physical constraints** | | | | | | |
| Hospitalized | 0.002 | 0.003 | 0.002 | 0.003 | 0.001 | 0.002 |
| Sick leave | 0.095 | 0.060 | 0.020 | 0.065 | 0.053 | 0.017 |
| Crime | 0.119 | 0.064 | 0.037 | 0.127 | 0.068 | 0.045 |
| Separation | 0.123 | 0.059 | 0.046 | 0.131 | 0.063 | 0.042 |
| Share with any physical constraints | 0.286 | 0.166 | 0.099 | 0.274 | 0.165 | 0.100 |
| Average number of physical constraints | 0.339 | 0.185 | 0.105 | 0.326 | 0.185 | 0.107 |
| **Norm constraints** | 0.602 | 0.480 | 0.267 | 0.583 | 0.485 | 0.315 |
| 80% men and low PL uptake at workplace | | | | | | |
| **Share with any constraint** | 0.931 | 0.748 | 0.721 | 0.910 | 0.737 | 0.742 |
| **Average number of constraints** | 1.928 | 1.200 | 1.128 | 1.869 | 1.197 | 1.242 |
| Observations | 6462 | 38070 | 5250 | 7002 | 37091 | 4375 |

*Note:* The table reports the prevalence of constraints among low-uptake fathers of children born in 1995–2015. Parental leave uptake refers to the number of paid parental leave days during the child's first two years. Low uptake refers to taking at most half of the reserved quota, which corresponds to 0–15 days for fathers of children born in 1995–2001 and 0–30 days for those of children born in 2002–2015. Pre-birth earnings are measured in the calendar year before childbirth and include annual labor income from wage- and self-employment as well as capital income. Definitions of the constraints are provided in Section 2.3 of the paper.

## Web Appendix W3. More detailed descriptive analysis of constraints that impact fathers' paid parental leave uptake by income group

Figures W6 through W8 offer a detailed breakdown of fathers' uptake of parental leave, depending on various observed constraints, across different earnings levels and over time. Figure W6 demonstrates that physical constraints have minimal explanatory power regarding income distribution.

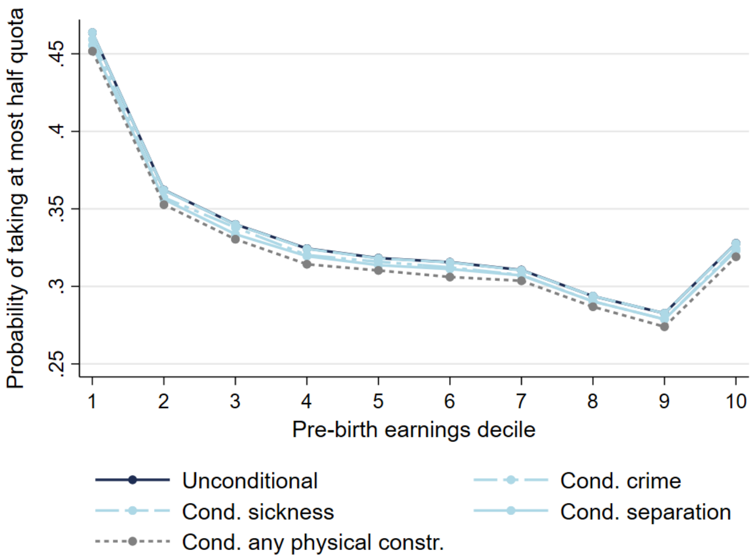


**Figure W6. The share of fathers by the degree of parental leave uptake and earnings group for children born unconditionally and conditional on physical constraints**
Note: The figure shows the share of fathers who take less than half the quota of the parental leave by pre-birth earnings decile unconditionally on any constraint (black solid line), conditional on sickness (blue small-dashed line), conditional on previously being convicted of a crime (blue long-dashed line), conditional on being separated (blue solid line), and conditional on any physical constraint (grey dashed line). The quota is 1-15 days from 1995–2001 and 1-30 days from 2002.

Figure W7 illustrates the evolution of low leave uptake conditional on economic constraints over time. For low- and high-income fathers, economic constraints—particularly being the primary earner and self-employment—are associated with persistently higher probabilities of low uptake, with little decline over time. In contrast, for middle-income fathers, these economic constraints appear less predictive, suggesting that other barriers dominate.

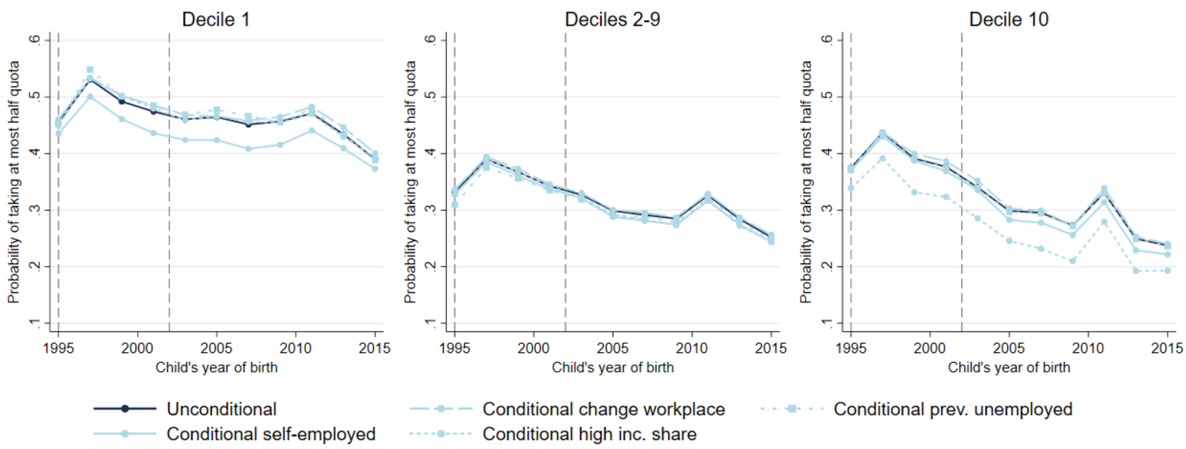


**Figure W7. The share of fathers by the degree of parental leave uptake and earnings group for children born unconditionally and conditional on economic constraints by year**
Note: The figure shows the share of fathers who take less than half the quota of the parental leave by pre-birth earnings decile unconditionally on any constraint (black solid line), conditional on changing workplace (blue dashed line), conditional on previously being unemployed (blue dotted line), conditional on being self-employed (blue solid line), conditional on having a high share of household income (blue short-dashed line), and conditional on any economic constraint (grey dashed line) and conditional on any constraint, be it physical, economic, or norms (grey dashed line). The quota is 1-15 days from 1995–2001 and 1-30 days from 2002.

Figure W8 complements this analysis by showing that physical constraints are relatively stable in their influence over time, particularly for low-income fathers.

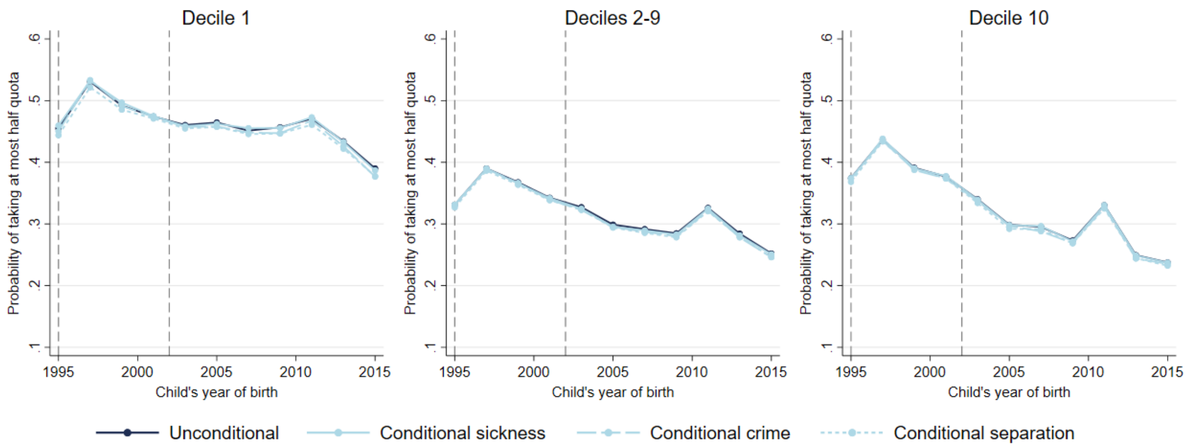


**Figure W8. The share of fathers by the degree of parental leave uptake and earnings group for children born unconditionally and conditional on physical constraints by year**
Note: The figure shows the share of fathers who take less than half the quota of the parental leave by pre-birth earnings decile unconditionally on any constraint (black solid line), conditional on sickness (blue small-dashed line), conditional on previously being convicted of a crime (blue long-dashed line), conditional on being separated (blue solid line), and conditional on any physical constraint (grey dashed line). The quota is 1-15 days from 1995–2001 and 1-30 days from 2002.

**Table W3. Percentage change of different constraints over income groups and time**

| | 1995 | 1999 | 2003 | 2007 | 2011 | 2015 |
|---|---|---|---|---|---|---|
| **Decile 1** | | | | | | |
| Conditional any | 15,82 | 7,11 | 22,78 | 23,5 | 21,66 | 23,33 |
| Conditional physical | 2,42 | 0 | 2,6 | 1,33 | 2,12 | 6,41 |
| Conditional economic | 7,69 | 4,27 | 8,03 | 6,43 | 5,31 | 2,31 |
| Conditional norms | 8,13 | 9,76 | 15,4 | 14,41 | 13,8 | 11,79 |
| **Deciles 2-9** | | | | | | |
| Conditional any | 16,92 | 16,03 | 20,49 | 23,97 | 23,01 | 32,14 |
| Conditional physical | 0,91 | 2,17 | 2,75 | 4,11 | 3,07 | 4,76 |
| Conditional economic | 8,76 | 4,08 | 6,12 | 7,88 | 7,36 | 7,14 |
| Conditional norms | 9,37 | 12,5 | 15,29 | 16,1 | 15,03 | 22,22 |
| **Decile 10** | | | | | | |
| Conditional any | 14,44 | 19,69 | 23,53 | 37,29 | 26,28 | 30,25 |
| Conditional physical | 1,87 | 1,53 | 2,94 | 3,39 | 2,72 | 3,36 |
| Conditional economic | 11,76 | 15,09 | 14,71 | 26,44 | 18,73 | 21,01 |
| Conditional norms | -0,27 | 5,88 | 8,82 | 12,2 | 9,37 | 12,18 |

## Web Appendix W4. More on the role of gender norms

We use regional data on gender attitudes from the World Value Survey, using data for the three years of available data for our study period, i.e., 1996, 2006, and 2011 (see Inglehart et al (2014) for the dataset and www.worldvaluessurvey.org for codebooks and detailed documentation). The WVS is nationally representative, and we use the county-level distribution of responses to three statements from the survey, each reflecting traditional beliefs: (i) 'When jobs are scarce, men should have more right to a job than women', (ii) 'University is more important for a boy than for a girl', and (iii) 'Men make better political leaders than women do'. For each statement, we compute the share of respondents who report agreement ("agree" or "strongly agree"), with higher values indicating more traditional norms. Non-responses and "don't know" answers are excluded. We exclude counties with fewer than 50 respondents per wave.

**Table W4. Descriptive statistics on gender attitude variables from the World Value Survey**

| Region | 'When jobs are scarce, men should have more right to a job than women' | 'University is more important for a boy than for a girl' | 'Men make better political leaders than women do' |
|---|---|---|---|
| Stockholm | 0.028 | 0.028 | 0.094 |
| Uppsala | 0.086 | 0.086 | 0.241 |
| Södermanland | 0.095 | 0.048 | 0.190 |
| Östergötland | 0.116 | 0.047 | 0.093 |
| Jönköping | 0.030 | 0.017 | 0.160 |
| Kronoberg | 0.077 | 0.103 | 0.205 |
| Kalmar | 0.091 | 0.091 | 0.091 |
| Blekinge | 0.172 | 0.034 | 0.069 |
| Skåne | 0.043 | 0.055 | 0.101 |
| Halland | 0.034 | 0.034 | 0.069 |
| Västra Götaland | 0.043 | 0.056 | 0.114 |
| Värmland | 0.167 | 0.104 | 0.292 |
| Örebro | 0.042 | 0.056 | 0.042 |
| Gävleborg | 0.045 | 0.136 | 0.273 |
| Västernorrland | 0.040 | 0.060 | 0.240 |
| Västerbotten | 0.083 | 0.104 | 0.250 |
| Norrbotten | 0.029 | 0.118 | 0.206 |
| Mean | 0.045 | 0.051 | 0.120 |
| Standard deviation | 0.035 | 0.039 | 0.063 |

*Note:* The table show the average share agreeing to the statements "When jobs are scarce men should have more right to a job than women" and "University is more important for a boy than for a girl", and "Men make better political leaders than women do", collected from the World Value Survey for the waves in 1996, 2006, and 2011. We include regions with at least 20 respondents per wave.

Table W4 presents descriptive statistics of the three regional proxies at the county level. While relatively few respondents agree with the gender norm, agreement is highest for the statement on political leadership, with an average of 12 percent and a county maximum of 29.2 percent.

For the other two statements, average agreement is 4.5 and 5.1 percent, respectively, again with substantial regional variation.

## Web Appendix W5. More on gender norms on the family level

To validate that having a younger sister is a valid proxy for gender norms, we estimate first-stage regressions of sibling-sex composition on gendered educational and occupational outcomes, following Brenøe (2021). These measures capture individuals' own gender-typical educational and occupational choices and serve as indicators of adherence to traditional gender norms among men. Specifically, we examine the following outcomes: (1) the average log male share in occupations held between ages 31 and 45; (2) the average share of years between ages 31 and 45 spent in male-dominated occupations; (3) the average share of years between ages 31 and 45 spent in STEM occupations; (4) the male share in the highest-attained educational field by age 30; (5) an indicator for whether the male share in the highest-attained educational field by age 30 is at least 80 percent; and (6) an indicator for whether the highest-attained educational field by age 30 is in a STEM discipline.

Table W5 presents the first-stage estimates on gendered occupational and educational outcomes. The results indicate that fathers with a younger sister, compared to those with a younger brother, spend more time in male-dominated occupations and are also more likely to pursue gender-typical education (Panel A). In Panel B, we examine whether these effects vary by grandparental education. This analysis is motivated by prior research suggesting that parents' educational attainment influences the transmission of gender norms to their children. Specifically, more educated parents may hold more egalitarian views and thus engage in less gender-stereotypical parenting (Geisler & Kreyenfeld, 2011). Applied to the sibling-sex composition, heterogeneity in grandparental education could attenuate the effect of the quasi-experimental variation in fathers' gendered upbringing. To account for this, we construct an indicator variable equal to one if the father has at least one parent with a university degree and no parent with only primary education, and zero otherwise. We interact this indicator with the dummy variable for having a younger sister.

**Table W5: Effect of having an opposite-sex sibling on fathers' probability to have gender-typical occupation or education**

| VARIABLES | (1) Log male share in occupation | (2) # years in male-dominated occupation | (3) # years in male-dominated occupation | (4) Log male share in education major | (5) Male-dominated education | (6) STEM education |
|---|---|---|---|---|---|---|
| *A. Average* | | | | | | |
| Second-born sister | 0.004 | 0.000 | 0.005** | 0.008** | 0.007* | 0.006 |
| | (0.003) | (0.003) | (0.002) | (0.004) | (0.004) | (0.003) |
| Observations | 78,170 | 77,992 | 78,170 | 79,954 | 79,954 | 79,954 |
| R-squared | 0.027 | 0.040 | 0.061 | 0.028 | 0.037 | 0.029 |
| *B. By grandparents' education* | | | | | | |
| Second-born sister | 0.004 | 0.000 | 0.007** | 0.012*** | 0.011** | 0.008* |
| | (0.004) | (0.004) | (0.003) | (0.004) | (0.004) | (0.004) |
| Second-born sister x High | -0.001 | -0.002 | -0.003 | -0.011* | -0.012* | -0.008 |
| | (0.006) | (0.007) | (0.005) | (0.006) | (0.007) | (0.006) |
| Observations | 78,170 | 77,992 | 78,170 | 79,954 | 79,954 | 79,954 |
| R-squared | 0.017 | 0.028 | 0.042 | 0.014 | 0.021 | 0.017 |

*Note:* The table shows the effect of having an opposite-sex sibling on fathers' gendered occupational and educational outcomes, including log male shares in occupations and education fields, years spent in male-dominated or STEM occupations, and indicators for male-dominated and STEM education majors. Full definitions are provided in the text. Panel A reports average effects, while Panel B shows estimates by grandparents' education. All models control for the father's birth county and year, sibling spacing, and grandparents' age and education. In columns (1)–(4), grandparental education is coded as level-by-field indicators; in columns (5)–(8), it is defined as an indicator equal to one if at least one grandparent has a university degree and no grandparent has only primary education ("Higher"). Robust standard errors, clustered at the birth county level, are reported in parentheses. *, **, and *** indicate statistical significance at the 10-, 5-, and 1-percent levels, respectively.

Differentiating by grandparents' education increases both the magnitude and precision of the estimates (see panel B). The results show that the estimated effects are generally positive and significant among fathers from less-educated families. In contrast, the interaction terms for fathers from highly educated families are negative in several outcomes, suggesting that the influence of sibling-sex composition on gendered occupational and educational choices is weaker—or may even reverse—among men raised in more highly educated families. This pattern aligns with the notion that parents with higher education may transmit more egalitarian norms and engage in less gender-stereotypical parenting. Overall, the findings from Panels A and B consistently support the validity of sibling-sex composition as a proxy for exposure to gender norms.[16]

**Table W6: Effect of having an opposite-sex sibling on fathers' probability to have gender-typical occupation or education, by birth cohort**

| VARIABLES | (1) Log male share in occupation | (2) # years in STEM occupation | (3) # years in male-dominated occupation | (4) Log male share in education major | (5) Male-dominated education | (6) STEM education |
|---|---|---|---|---|---|---|
| *A. 1995–2000* | | | | | | |
| Second-born sister | 0.007 | -0.005 | 0.010** | 0.011** | 0.010** | 0.006 |
| | (0.006) | (0.005) | (0.004) | (0.005) | (0.004) | (0.004) |
| Observations | 25,075 | 24,918 | 25,075 | 24,545 | 24,545 | 24,545 |
| R-squared | 0.025 | 0.044 | 0.058 | 0.031 | 0.039 | 0.030 |
| *B. 2001–2005* | | | | | | |
| Second-born sister | 0.003 | 0.003 | 0.002 | 0.005 | 0.004 | 0.005 |
| | (0.004) | (0.003) | (0.004) | (0.006) | (0.006) | (0.005) |
| Observations | 35,305 | 35,286 | 35,305 | 35,421 | 35,421 | 35,421 |
| R-squared | 0.028 | 0.047 | 0.069 | 0.028 | 0.037 | 0.031 |
| *C. 2010–2015* | | | | | | |
| Second-born sister | 0.003 | 0.000 | 0.006 | 0.008 | 0.010 | 0.005 |
| | (0.007) | (0.003) | (0.007) | (0.008) | (0.008) | (0.007) |
| Observations | 17,790 | 77,992 | 17,790 | 19,988 | 19,988 | 19,988 |
| R-squared | 0.031 | 0.021 | 0.062 | 0.027 | 0.032 | 0.033 |

*Note:* The table shows the effect of having an opposite-sex sibling on fathers' gendered occupational and educational outcomes, including log male shares in occupations and education fields, years spent in male-dominated or STEM occupations, and indicators for male-dominated and STEM education majors. Full definitions are provided in the text. Panel A reports average effects, while Panel B shows estimates by grandparents' education. All models control for the father's birth county and year, sibling spacing, and grandparents' age and education. In columns (1)–(4), grandparental education is coded as level-by-field indicators; in columns (5)–(8), it is defined as an indicator equal to one if at least one grandparent has a university degree and no grandparent has only primary education ("Higher"). Robust standard errors, clustered at the birth county level, are reported in parentheses. *, **, and *** indicate statistical significance at the 10-, 5-, and 1-percent levels, respectively.

In addition to our robustness checks for the second-stage outcome, we examine whether the relationship between sibling-sex composition and gendered occupational and educational choices—the first-stage outcomes in our framework—varies across cohorts. Table W6 presents

[16] To assess if having a younger sister is a valid proxy for traditional gender norms for all childbirth cohorts, Web Appendix Table W7 presents estimates form separate regression for each cohort, e.g., for fathers who had a child in 1995–2001, 2002–2009, and 2010–2015. While precision is lower, the estimates consistently point in the same direction as those in Table 4 for all cohorts.

estimates for fathers born in three periods: 1995–2000, 2001–2005, and 2010–2015. The results indicate that the first-stage associations are generally positive and consistent in direction across cohorts, although the magnitudes and statistical significance vary somewhat. For instance, fathers with a younger sister are significantly more likely to spend time in male-dominated occupations and to pursue gender-typical education in the earliest cohort (1995–2000). In the subsequent cohorts, the estimates remain positive but are smaller in magnitude and generally not statistically significant.

These findings suggest that while the strength of the association may vary somewhat across birth cohorts, there is no clear evidence of systematic trends over time in the relationship between sibling-sex composition and gender-typical occupational or educational outcomes. This supports the stability of sibling-sex composition as a proxy for exposure to gender norms across different generations of fathers.

We conduct several robustness analyses to assess the sensitivity of our main results regarding the effect of sibling-sex composition on fathers' parental leave uptake. First, in Panel A of Table W7, we relax the restriction on family size and include all fathers with at least two siblings, rather than restricting the sample to two-child families. The estimates remain small and statistically insignificant in most cohorts, except for a modest positive effect in the 2010–2015 cohort, where having a younger sister increases the probability of taking at most half of the reserved leave days by approximately 0.9 percentage points. Second, in Panel B of Table W7, we further include a control for the total number of siblings to account for potential confounding from family size. The results are virtually unchanged compared to the unrestricted sample without this control, suggesting that our findings are not driven by differences in family size.

**Table W7: Effect of having a younger sister on fathers' probability of taking at most half of reserved parental leave days – robustness analyses**

| VARIABLES | (1) 1995-2015 | (2) 1995-2001 | (3) 2002-2009 | (4) 2010-2015 | (5) 1995-2015 | (6) 1995-2001 | (7) 2002-2009 | (8) 2010-2015 |
|---|---|---|---|---|---|---|---|---|
| *A. No restriction of family size* | | | | | | | | |
| Second-born sister | 0.002 | 0.003 | -0.002 | 0.007 | 0.003 | 0.003 | -0.000 | 0.009* |
| | (0.003) | (0.004) | (0.005) | (0.004) | (0.003) | (0.005) | (0.005) | (0.004) |
| Second-born sister x Higher grandparental education | | | | | -0.002 | 0.003 | -0.004 | -0.006 |
| | | | | | (0.003) | (0.009) | (0.007) | (0.007) |
| Observations | 144,807 | 41,177 | 58,987 | 44,643 | 144,807 | 41,177 | 58,987 | 44,643 |
| R-squared | 0.022 | 0.019 | 0.028 | 0.038 | 0.019 | 0.017 | 0.025 | 0.038 |
| *B. No restriction of family size + control for number of siblings* | | | | | | | | |
| Second-born sister | 0.002 | 0.004 | -0.002 | 0.006 | 0.003 | 0.003 | -0.000 | 0.008* |
| | (0.003) | (0.005) | (0.005) | (0.004) | (0.003) | (0.005) | (0.005) | (0.004) |
| Second-born sister x Higher grandparental education | | | | | -0.002 | 0.003 | -0.004 | -0.006 |
| | | | | | (0.003) | (0.009) | (0.008) | (0.007) |
| Observations | 144,807 | 41,177 | 58,987 | 44,643 | 144,807 | 41,177 | 58,987 | 44,643 |
| R-squared | 0.022 | 0.019 | 0.028 | 0.038 | 0.019 | 0.017 | 0.025 | 0.033 |
| *C. Control for child gender* | | | | | | | | |
| Second-born sister | 0.002 | -0.002 | -0.001 | 0.012** | 0.002 | -0.000 | -0.002 | 0.012** |
| | (0.003) | (0.005) | (0.005) | (0.004) | (0.003) | (0.005) | (0.006) | (0.006) |
| Second-born sister x Higher parental education | | | | | 0.001 | -0.005 | 0.004 | -0.002 |
| | | | | | (0.005) | (0.011) | (0.008) | (0.011) |
| Observations | 87,742 | 25,235 | 36,038 | 26,469 | 87,742 | 25,235 | 36,038 | 26,469 |
| R-squared | 0.022 | 0.023 | 0.028 | 0.040 | 0.019 | 0.020 | 0.024 | 0.036 |

*Notes:* The table shows the effect of having an opposite-sex sibling on fathers' probability of taking at most half of the reserved parental leave days during the child's first two years. Taking at most half of the reserved days corresponds to a maximum of 15 days for children born before 2002 and 30 days for those born in 2002 or later. In panel A and B, the sample includes fathers with two or more siblings. All models control for the father's county and year of birth, spacing to the younger sibling, and grandparents' age at the father's birth and their level of education. In columns (1)–(4), grandparental education is included as indicators for level-by-field of education. In columns (5)–(8), grandparental education is defined as an indicator equal to one if at least one grandparent has a university degree and no grandparent has only primary education ("Higher"). Robust standard errors, clustered at the birth county level, are reported in parentheses. *, **, and *** indicate statistical significance at the 10-, 5-, and 1-percent levels, respectively.

Third, in Panel C, we return to our main sample of two-child families and add a control for the gender of the child for whom the parental leave is taken. This analysis is motivated by prior research suggesting that fathers' labor market outcomes and parental leave-taking behaviors may differ depending on the child's gender (e.g., Dahl & Moretti, 2008; Lundberg, 2005). Social expectations around father involvement might vary for sons versus daughters, potentially influencing leave decisions. If such mechanisms affect fathers' parental leave behavior, failing to account for child gender could bias our estimates of the effect of sibling-sex composition. Including this control does not substantively alter the results. The effect of having a younger sister remains positive and statistically significant for the 2010–2015 cohort, with an estimated increase of approximately 1.2 percentage points in the likelihood of taking at most half of the reserved leave days. Across other cohorts, the estimates remain small and statistically insignificant. Overall, these robustness checks confirm that our main findings are not sensitive to alternative sample restrictions or the inclusion of additional controls.

## References in the Web Appendix

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