Competition and Career Advancement: The Hidden Costs of Paid Leave

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Working Paper

2020-059

08/2020

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This version: July 2020

Abstract

Does leave-taking matter for young workers' careers? If so, why? We propose the competition effect—relative leave status of workers affecting their relative standing inside the firm—as a new explanation. Exploiting a policy reform that exogenously assigned four-week paid paternity leave to some new fathers, we find evidence consistent with the competition effect: A worker enjoys a better post-child earnings trajectory when a larger share of his colleagues take leave because of the policy. In contrast, we find no direct earnings effect resulting from the worker's own leave when controlling for their relative leave eligibility status within the firm.

Key words: leave of absence, career interruptions, ranking, tournament, promotion, gender gap

JEL codes: M51, M52, J16, J22, J24, J31

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* We thank Jerome Adda, Christian Dustmann, Henry Farber, Claudia Goldin, Ed Lazear, Jin Li, Katrine Løken, Claudia Olivetti, Jessica Pan, Imran Rasul, Uta Schoenberg, Kathryn Shaw, Linh Tô and seminar/conference participants at AASLE, NBER Summer Institute, HKU, NUS, NHH, UCL, and VATT-GSE Helsinki for helpful discussions and comments.

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"Showing up is 80 percent of life" (Woody Allen).

1. Introduction

In advanced economies, workers are entitled to various statutory leave benefits, including maternity, paternity, family, and sick leave. Yet despite guarantees of job protection and income replacement for the duration of the leave, workers often fear that taking it may harm their career paths and future earnings. In this paper, we investigate whether and why absence from work—even if brief initially—may affect the leave-taker's future earnings, by exploiting administrative data and the introduction of a four-week fully paid paternity leave in Norway.[1]

Investigating the possible career costs of leave, the literature has mainly focused on leave-engendered loss of human capital and productive skills (Mincer and Polachek 1974; Mincer and Ofek 1982; Adda et al. 2017), which we label as direct effect.[2] We instead put forward a different perspective, viewing the relative standing of workers inside the firm as a potentially important driver of worker outcomes, which we label as competition effect. For instance, if there are (pre-determined) high vs. low wage slots in the firm and the assignment to these slots is determined by the relative output or performance of the contestants (Lazear and Rosen 1981; Holmstrom 1982; Gibbons and Waldman 1999), then leave-taking may lower the chances of winning even in the absence of absolute skill loss. This will be the case if, for any number of reasons, an ever-present contestant is ranked higher than another contestant who is absent for part of the time, even when the two are ex ante identically skilled. For instance, a worker—while on leave or in anticipation of an upcoming leave—may miss out on assignment to high impact projects or tasks in the firm.

[1] Norway was the first European country to adopt, as early as 1993, a fully paid parental leave specifically reserved for fathers.

[2] The direct effect will likely be picked up in earnings if workers are paid by their absolute productivity as in performance pay (Lazear 2000; Lemieux et al. 2009).

Moreover, having fewer face-to-face interactions with colleagues or supervisors may hinder the transmission of tacit knowledge in workplaces (Sandvik et al. 2020), thereby affecting a worker's relative performance evaluation.[3]

Causally identifying the direct and competition effects in observational data—when all workers are leave eligible—is challenging, primarily because workers self-sort into leave-takers and non-takers. For instance, less productive workers may select into more frequent or longer leave periods than others because of their low earnings and the low opportunity cost of leave (Becker 1985; Polachek 1981; Gronau 1988; Adda et al. 2017).[4] Likewise, when information asymmetry is prevalent, workers may try to signal their types by choosing different decisions with respect to leave (Tô 2018).[5] In both scenarios, the lower observed earnings of leave-takers relative to non-takers is a reflection of their underlying types, rather than a causal effect of exogenously assigned leave for comparable workers.[6] Moreover, even if the sorting problem is resolvable—e.g. through exogenous assignment of workers to leave-takers and non-takers—any credible research design that focuses on the effect of a worker's *own* leave on his own earnings is likely to identify a combination of the direct and competition effects, which we label the total effect.[7] In this paper

[3] Further, there are many anecdotal examples suggesting gains to being present and visible in the firm as opposed to being absent. For instance, Bloom et al. (2014) show that working from home disadvantages workers on promotion chances even when their productivity is not lowered. Also, Armenti (2004) shows that female assistant professors often time their pregnancy such that childbirth occurs in May, just when the summer break starts, to make their maternity leave least salient.
[4] Albanesi and Olivetti (2009) point out that gender differences in the opportunity cost of market (vs. home) work may emerge in a self-fulfilling equilibrium even in the absence of exogenous gender differences in productivity.
[5] Albrecht et al. (1999) also invoke a signaling story when discussing their findings from Sweden that different types of time out (e.g. parental leave, household time, unemployment) have different effects on wages, which vary by gender and are unexplainable by human capital depreciation alone. Also, Thomas (2018) illustrates a case where the provision of a paid leave policy triggers statistical discrimination by firms due to changing composition of women in the workforce induced by the policy.
[6] Note that in signaling models, *correlation* (vs. causation) between types, actions, and outcomes is precisely what the theory predicts and what researchers should look for.
[7] That is, if a worker takes leave, his skills or human capital may depreciate from the state of not working. Moreover, the fact that he is away (and hence misses out on promotable tasks and opportunities) may lower his ranking against his peers even in the absence of absolute skill loss. Therefore, a comparison of (exogenous) leave-taker with (exogenous) non-taker must identify a combination of the skill loss and competition effects.

we address both of these challenges, and disentangle the competition and direct effects, highlighting the significance of the former.

Our empirical analysis takes advantage of a 1993 policy reform in Norway that provides four weeks of paid paternity leave for fathers of children born on or after April 1, 1993, with full pay furnished by the social insurance system and not the individual firm.[8] Most important, these four weeks are reserved for fathers only, with no unused portions transferable to the child's mother.[9] The outcome of this reform was a sudden and dramatic increase in the share of paternity leave taken by fathers of children born after the policy cutoff date (see Figure 1).[10]

To isolate the hypothesized competition effect, we propose a novel identification strategy that exogenously shifts contestants (i.e. a set of focal fathers and their competitors inside the firm) up and down the ranking, not via the focal father's own leave but through the reform-induced leave of his competitors. The focal fathers are defined as male workers who had a first child during a narrow window spanning 9 months before and after the policy cutoff date. For each focal father, we identify a set of relevant competitors, defined as male workers in the same plant—hereafter referred to as firm—who also fathered a child in the 18 month window and have similar age and education as the focal father. We then calculate the share of each focal father's competitors who are leave eligible (i.e. whose children are born after the policy cutoff date), which is arguably exogenous to the focal father's own eligibility for the paid paternity leave but correlates with the

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[8] See Kotsadam and Finseraas (2011), Rege and Solli (2013), Dahl et al. (2014), and Cools et al. (2015), for prior research exploiting this policy reform.

[9] Prior to the 1993 reform, if the father did not use any of the *shared* parental leave, then the mother could use it, so the father's failure to take paternity leave did not imply a wasted benefit at the household level. The newly introduced four-week paternity quota is in addition to the existing amount of shared leave, which the couple can divide any way they wish both before and after the 1993 reform.

[10] As to the length of leave taken, Figure A1 shows clear bunching of leave duration around four weeks, as specified by the paid paternity leave quota.

share of his competitors taking leave.[11] By comparing the career trajectories (-3 to 7 years since child birth) of focal fathers exposed to a high versus low share of leave-eligible (vs. ineligible) competitors, we identify the competition effect driven by the asymmetric leave status of the contestants in a difference-in-differences (DID) framework.

We find that—holding fixed the focal father's status at either policy ineligible or eligible—when a larger share of his competitors are eligible for the paid paternity leave, the focal father's own post-child earnings are put on a higher trajectory than would otherwise have been the case. In particular, our reduced-form (intention-to-treat) estimates show that when the share of births to leave eligible (vs. ineligible) competitors increases by 0.37 (1 *SD* in the sample)—which is equal to an additional two out of five competitor births (the sample average)—a focal father's earnings are on average 1.1 percent higher annually in the 7 years following the child's birth than they would have been otherwise. This competition effect is concentrated in firms where the earnings growth of workers is more dispersed and the competition aspect is likely to be more relevant in determining their earnings.[12]

We then investigate the total effect of the focal father's own leave on his earnings. Here, unlike the competitors' leave, which elevates only the focal father's relative position, the focal father's leave induces both a direct and competition effect. Our reduced-form (intention-to-treat) estimates reveal a negative total effect of focal father leave eligibility that is close in magnitude to the competition effect established above. However, once we shut down the competition channel (by conditioning on the difference between the leave eligibility of the focal father and that of his competitors), the effect of the focal father's own leave on his earnings disappears, suggesting that

[11] As Figure 2 illustrates, the share of competitors *eligible* for leave has strong predictive power for the share *actually taking* leave, which is the basis of our IV analysis.
[12] On the contrary, if the pay scale is pre-determined for workers of the same tenure and skill levels, for instance, the relative absence/presence of the contestants would have little impact on their earnings.

the direct effect from absolute skill loss is unlikely to be the main driver of the earnings penalty associated with leave.

Overall, the results we derive in this context of policy-driven paternity leave indicate that temporary absence matters for young workers' careers, not because of the skill depreciation it may cause but because (exogenous) leave-takers lose out to (exogenous) nontakers. These initial leave effects may be augmented if leave-taking competitors remain more attached to their families (vs. the employer) than non-taking competitors, even after the initial leave has ended.[13] The competition effect documented here should thus be interpreted as the reduced-form effect of the policy-driven four-week paid paternity leave, inclusive of any subsequent modifications in worker behavior that the initial leave triggers.[14]

A major contribution of our analysis is that by cleanly identifying the competition effect of leave in male-to-male contests, it isolates an important yet previously unexplored channel through which absence from work (due to parental leave) affects young workers' careers. Even in European countries that provide relatively generous parental leave, the system has been such that female workers by default take a long maternity leave after childbirth whereas their male colleagues do not (see e.g. Lalive and Zweimüller 2009; Schönberg and Ludsteck 2014).[15] In such situations, the logic of the competition effect would imply that policies that shut down the between-worker asymmetry in leave positions and thereby make the competition channel of leave inoperative—e.g. universal paternity leave or normalization of remote working—should facilitate

[13] For instance, Patnaik (2019) and Farré and González (2019) show that, in response to five weeks of paternity leave in Quebec and two weeks in Spain, respectively, leave eligible fathers substantially modify their behavior, spending more time on non-market work including childcare. In contrast, Ekberg et al. (2013) find no evidence of increased participation in household work among men treated by the "daddy month" reform in Sweden.

[14] When considering the policy impact of paternity leave quota, this reduced-form effect is the relevant parameter.

[15] Interestingly, Kleven et al. (2019) show that the motherhood penalty in terms of labor earnings is larger in countries with longer maternity leave such as Austria and Germany.

equal career progression of male and female workers after childbirth.[16] There is a growing body of work that identify occupational/workplace characteristics and/or remuneration policies that hinder equal advancement of men and women on the firm or professional hierarchy, including Lazear and Rosen (1990), Landers et al. (1996), Black and Strahan (2001), Bertrand et al. (2010), Goldin (2014), Card et al. (2016), Goldin and Katz (2016), Azmat and Ferrer (2017), Babcock et al. (2017), Bagues et al. (2017), Barth et al. (2017), and Sarsons (2019) among others.[17] We add to this debate by pointing out the competition effect of leave in workplaces and highlighting the value of being present (when others are).

2. Paternity Leave Reform and Theoretical Framework

2.1 Paternity leave reform of 1993

The universal Norwegian parental leave scheme, which is part of the broader Social Security System financed through worker and firm taxes, offers not only employment protection but also a 100% earnings replacement. While the replacement is capped, this cap is nonbinding for most parents because all public and most private employers compensate for any earnings above it. Prior to the 1993 reform (announced in October 1992 and passed in parliament in December 1992), the paid parental leave period comprised 35 weeks, eight of which were reserved for the mother. The parents could then share the remaining 27 weeks between them as desired, although only three percent of fathers actually took any leave (see Figure 1).

[16] This will be the case only if such gender-neutral policies have the same ramifications for the productivity of male and female workers. Antecol et al. (2018) present a situation where male academics use parental leave for producing more research, which may hurt their female counterparts who cannot do the same.

[17] See Altonji and Blank (1999), Bertrand (2011), and Blau and Kahn (2017) for a broader discussion of factors contributing to gender earnings gap.

Because of this low uptake, on April 1, 1993, the Norwegian government enacted a paternity leave policy whose explicit aim was to promote gender equality in the use of paid parental leave. This reform expanded the paid parental leave period to 42 weeks, with four weeks reserved specifically for the father and nontransferable to the mother, meaning that they are lost to the family if the father does not use them.[18] The policy was announced only 6 months before its implementation, which means that fertility decisions for births around the implementation date were already taken before announcement and could not have been affected by the policy itself.

2.2 Theoretical framework and implications of the reform

We present a simple stylized model of a two-player contest. The contestants are two workers, the focal father and the competitor, with identical abilities and preferences who just fathered a child. Each player's objective is to maximize individual utility, which comprises utility at home and wages at work. Each contestant chooses whether or not to take paternity leave to maximize his individual utility. In this model, paternity leave affects both components of the individual utility, and equilibrium is a pair of actions chosen by the two players from which neither has an incentive to deviate.

We employ this model to illustrate how the 1993 paternity leave reform shifts the equilibrium in the contest modeled. More specifically, depending on the birthdate of the worker's child relative to the policy cutoff date (April 1, 1993), the contest may fall into four different game scenarios: *focal father ineligible/competitor ineligible*, *ineligible/eligible*, *eligible/ineligible,* and *eligible/eligible*, where eligible (vs. ineligible) refers to whether the worker's child birth falls after (vs. before) the implementation of the new policy.

[18] The leave reserved for the mother also expanded from eight to nine weeks.

Utility at home. We normalize the shared parental leave at unity, and designate the four weeks of paternity leave specified in the new policy by $\rho \in [0,1]$. For a father who had a child prior to the policy cutoff date (and is thus ineligible for the new paternity leave), utility at home is

$$B^N = \begin{cases} \rho b^f + (1-\rho) b^m & \textit{if taking paternity leave,} \\ b^m & \textit{if not taking paternity leave,} \end{cases}$$

where b^f is the nonpecuniary benefit (e.g. child welfare) when the father spends one unit of time with the child, and b^m is that when the mother does so. For simplicity, we assume that $b^m > b^f$.[19] Under the status quo, if the father takes leave for duration ρ, it eats into the leave available to the mother.

For a worker who has a child after the policy cutoff date (and is thus eligible for the paid paternity leave), utility at home is

$$B^Y = \begin{cases} \rho b^f + b^m & \textit{if taking paternity leave,} \\ b^m & \textit{if not taking paternity leave.} \end{cases}$$

Since the new paternity leave of ρ is additional to the existing shared leave (of length 1), the father's leave-taking no longer reduces the leave available to the mother. For simplicity, from here on, we set $\rho = 1$.

Wages at work. We interpret the competition effect through the lens of Lazear and Rosen's (1981) classic rank-order tournament, although the competition effect we put forward here is more generic and can be interpreted in many different ways. Each worker's choice variable is binary (to take or not take paternity leave). Since the two workers are otherwise identical, their relative leave status alone determines their rank in this model. The higher ranked worker receives W_H compared to

[19] Although not crucial for deriving the qualitative implications of the new paternity leave policy, this assumption ensures that no father takes paternity leave prior to policy introduction, which not only simplifies our analysis but is consistent with the empirical evidence of less than 3 percent of men taking leave before the 1993 reform.

$W_L (< W_H)$ for the lower rank, with an assumption of $W_H - W_L < 2b^f$.[20] In the case of tie, each worker receives $\frac{1}{2}W_H + \frac{1}{2}W_L \equiv W$. Because these wage levels are fixed ex ante, the employer's total wage bill remains fixed at $W_H + W_L$ regardless of which worker attains the higher wage of W_H ex post.

Worker payoffs. Each worker's payoff is the sum of the utility at home and the wages received at work. Depending on his child's birthdate, each worker can be either ineligible or eligible for the new policy. Below we analyze the payoffs of the focal father according to his policy eligibility status and under different contingencies with respect to his own and his competitor's leave choices.

If the focal father is policy ineligible, then his payoff is equal to:

$$(1) \quad V^N = \begin{cases} b^m + W & if\ no\ leave, no\ leave, \\ b^f + W_L & if\ leave, no\ leave, \\ b^m + W_H & if\ no\ leave, leave, \\ b^f + W & if\ leave, leave, \end{cases}$$

where the first and second actions in each row refer to that of the focal father and the competitor, respectively. Based on (1), we see that when the focal father is policy ineligible, not taking leave is his dominant strategy.[21] The intuition is that utility at home attainable from taking leave is lower than that from not taking leave—irrespective of the competitor's choices—and taking leave can only (weakly) lower his rank in the firm.

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[20] That is, the potential loss in wages due to leave is never too large. This assumption ensures a unique equilibrium in each subsequent scenario, which greatly simplifies our illustration of the policy's impact.

[21] If the competitor does not take leave, the focal father receives a higher payoff by not taking leave. Comparing rows 1 and 2 in (1), we can see that $b^m + W > b^f + W_L$, because $W - W_L = \frac{1}{2}(W_H - W_L) > 0$ and $b^m - b^f > 0$. If the competitor does take leave, the focal father again receives a higher payoff by not taking leave. From rows 3 and 4 in (1), we see that $b^m + W_H > b^f + W$, because $W_H - W = \frac{1}{2}(W_H - W_L) > 0$ and $b^m - b^f > 0$.

If the focal father is policy eligible, then his payoff is equal to:

$$(2)\quad V^Y = \begin{cases} b^m + W & \textit{if no leave, no leave,} \\ b^f + b^m + W_L & \textit{if leave, no leave,} \\ b^m + W_H & \textit{if no leave, leave,} \\ b^f + b^m + W & \textit{if leave, leave,} \end{cases}$$

where the first and second actions in each row again refer to that of the focal father and the competitor, respectively. When he is policy eligible, taking leave is the dominant strategy for the focal father, because the utility at home attainable from taking leave is higher than that from not taking it, and the assumption that $W_H - W_L < 2b^f$ guarantees that the potential loss in wages is never too large. [22]

Competitor's leave eligibility and equilibrium shifts. Based on the analysis above, we know that for the focal father—and also for the competitor since they are symmetric—not taking (taking) leave is the dominant strategy if he is ineligible (eligible) for the 1993 policy. Therefore, depending on the two workers' respective child birthdates (and eligibility for the 1993 policy), the equilibrium of this model will look as follows:

$$(3)\quad (\text{focal father, competitor}) = \begin{cases} (\textit{no leave, no leave}) & \textit{if ineligible/ineligible,} \\ (\textit{no leave, leave}) & \textit{if ineligible/eligible,} \\ (\textit{leave, no leave}) & \textit{if eligible/ineligible,} \\ (\textit{leave, leave}) & \textit{if eligible/ineligible,} \end{cases}$$

where the wages at work received by the focal father will be W, W_H, W_L and W, respectively (see the wage part of the payoffs in (1) and (2)).

[22] If the competitor does not take leave, the focal father receives a higher payoff from taking leave. From rows 1 and 2 in (2) and the assumption $W_H - W_L < 2b^f$, it follows that $b^f + b^m + W_L > b^m + W$. Similarly, when the competitor does take leave, the focal father obtains a higher payoff from taking leave: $b^f + b^m + W > b^m + W_H$ (rows 3 and 4 in (2)).

We then engage in a thought experiment in which the focal father's status is fixed at leave ineligible but that of his competitor is exogenously shifted from leave ineligible to eligible (i.e. from row 1 to row 2 in (3)). This exogenous shift from leave ineligible to eligible competitor status increases the focal father's wage at work from W to W_H. Likewise, in a second experiment, fixing the focal father's status at leave eligible and shifting his competitor's status from leave ineligible to eligible (i.e. from row 3 to row 4 in (3)) increases the focal father's wage at work from W_L to W. This prediction serves as the theoretical basis of our identification strategy in Section 3.

3. Empirical Framework

3.1 Sample

We build our dataset by combining data from several administrative registers, which cover the full Norwegian population and are linked through unique individual identifiers. The birth register not only gives the baby's expected birthdate and actual birth month but also links each individual birth to a mother and father identifier. Social security registers provide information on take-up of parental leave from 1992 onwards. Matched employer-employee registers provide data on firms and their employees, allowing us to match each worker to all his colleagues in the firm. We draw gross earnings data from the tax registers and deflate earnings to 1993. The national population register and education registers further provide annual information on marital/cohabitation status and educational attainment.

We restrict our sample to men who fathered their first child in the 18 months around the reform cutoff date (i.e. July 1, 1992 to December 31, 1993), with a history of stable employment in the same firm, defined as working there in the year of the child's birth and the two years prior. We drop focal fathers who had twins or multiple births in the 18-month observation period. We also restrict our sample to focal fathers aged—at child's birth—within the 10th and the 90th percentiles (25 and 36 years) of the age distribution to make focal fathers most comparable.[23] To investigate the focal fathers' career development, we track their log earnings from 3 years prior to 7 years after the birth of their first-born child. We therefore restrict the sample to those with a minimum level of labor market attachment by requiring that earnings be above one social security basic unit in each period.[24] For the pre-birth years, this criterion also proxies individual eligibility for paid paternity leave (Dahl et al. 2014).

We define each focal father's competitors as male co-workers who also fathered a child in the 18 month window and worked in the same firm in the year of the focal father's child birth and the two years prior. To capture only relevant competitors, we further define them as workers of similar age (±4 years) to the focal father and having the same educational level (2 years of high school or below; high school diploma; some college including a college degree; or beyond a college degree). To enable our analysis of competition effects, we restrict the sample to focal fathers with at least one competitor.[25] As this requires that at least two male workers (the focal father and one or more competitors) became a father during this window, our sample by construction excludes very small firms or firms with predominantly female employees.

Our focus on relatively large firms contrasts with Dahl et al. (2014)—a well-known research based on the same 1993 reform as used here—whose sample is by construction restricted to relatively small firms in which only one male employee has a childbirth within a narrow window around the policy cutoff date. The child birth date of that unique employee (i.e. ineligible or

[23] For robustness, we also estimate our main specifications without this age restriction.
[24] In 1993, one social security basic unit (1G) corresponds to about 18% of the average earnings in Norway, and the average earner would have to work around 6.6 hours per week to earn 1G.
[25] We also drop focal fathers and competitors in instances when at least one competitor had twins or multiple children within the 18-month observation window.

eligible for the paid paternity leave quota) provides the key variation in Dahl et al. (2014) in their analysis of leave take-up by the unique employee's co-workers having a child sometime in the future (between 1993 and 2006). In contrast, both the focal fathers and their competitors in our analysis have their children's births within a narrow window (1992-1993) straddling the policy cutoff date, which we use to exogenously vary the composition of their child birth dates (pre or post policy cutoff date) and thereby the contestants' relative leave status.

In Table 1, we report the baseline characteristics of the focal fathers in our sample as measured in the year of their child's birth, first for the overall sample, and then separated by high versus low share of competitor births occurring after the policy cutoff date. On average, a focal father has 4.63 births competitors fathering a child during the same window. As the table shows, the focal fathers' baseline characteristics are highly balanced in the two cases, confirming that the share of competitors eligible for the paid paternity leave is largely exogenous to the focal fathers.

3.2 Identification strategy

We isolate the paternity leave's competition effect using exogenous variation in the competitors' leave eligibility status while holding focal father's own status fixed at either leave ineligible or eligible. This identification strategy is analogous to exogenously shifting the contest scenario from *ineligible/ineligible* to *ineligible/eligible* (from row 1 to row 2 in (3)) or from *eligible/ineligible* to *eligible/eligible* (from row 3 to row 4 in (3)). Empirically, we achieve this goal by exploiting the child's birthdate composition (pre or post policy cutoff date) among each focal father's competitors.

Each focal father has at least one competitor having a child within our observation window. Crucially, the share of competitors' births falling within the latter half of this window is (arguably) exogenous to the focal father.[26] Hence, by comparing the career trajectories (-3 to 7 years since the child's birth) of focal fathers exposed to a high versus low share of leave-eligible competitors (among all competitors fathering a child within the window), we can identify the competition effect on the focal fathers in a difference-in-differences (DID) framework.

Specifically, in a sample of focal fathers, we estimate the following equation:

$$(4) \quad y_{it} = \phi_i + \psi_{\tau(it)} + \sum_{\tau=-3}^{7} \beta_\tau I(it = \tau) \times C_i + \boldsymbol{G}_{it}\lambda + u_{it},$$

where y_{it} is the log earnings of focal father i in year t. The event year, or year-since-birth (YSB), is denoted by τ, with the child's birth year designated as zero. The variable C_i measures the ratio of i's competitors that take paternity leave (conditional on having a child) to the total number of i's competitors having a child. We define $\tau = -1$ as the reference category such that all β_τs are relative to the year before the focal father's child birth. Vector $\boldsymbol{G}_{it}$ includes additional controls that vary at the worker-year level (e.g. a cubic polynomial in age). We cluster standard errors at the initial firm level, within which the competition group for each focal father is defined.

Conditioning on individual fixed effects (ϕ_i) and YSB fixed effects ($\psi_{\tau(it)}$), we estimate the evolution of the focal father's earnings after his child's birth dependent on his competitors' leave status, relative to the earnings in $\tau = -1$ (the year before the child birth). The Equation (4) corresponds to an event study that allows comparison of the pre-trends prior to the child's birth.

We also estimate a variant of (4) that collapses the β_τs for all the post-birth periods and estimates a single parameter β:

$$(5) \quad y_{it} = \phi_i + \psi_{\tau(it)} + \beta Post_{\tau(it)} \times C_i + \boldsymbol{G}_{it}\lambda + u_{it},$$

[26] Not only are the observable characteristics of focal fathers with high versus low shares of leave eligible competitors balanced (Table 1), but the pre-birth trends of log earnings between them are also comparable (Figure 3).

where $Post_{\tau(it)}$ indicates whether the period is in the year of or after the child's birth ($\tau \geq 0$). We estimate equation (5) first by OLS and then by IV, with C_i (share of competitors taking the paternity leave) instrumented by S_i (share of competitors eligible for it). As Figure 2 shows, S_i has strong predictive power for C_i, which is anticipated in Figure 1 by the discrete jump around the policy cutoff date. The OLS and IV estimates differ in two key ways: (i) whereas OLS uses the variation in C_i, which is driven by both choice (by competitors) and policy, IV uses only the policy-driven variation in C_i; and (ii) OLS is concerned with the average focal father/contest in the sample, whereas IV is based on the marginal focal father/contest whose C_i is shifted because of S_i.

Our research design fixes the focal father's own eligibility status while exogenously varying those of his competitors. It thus induces only the competition effect (or rank effect) free of any direct effect, unlike designs that exploit an exogenous shift in the focal father's own eligibility status and identifies the total—absolute skill loss plus competition—effect (see Section 4.3).

4. Career Costs of Paternity Leave

4.1 Competition effect

Figure 3 presents a first visual impression of the competition effect. The figure illustrates the raw trends in log earnings of focal fathers for whom the share of competitors eligible for the paid paternity leave is more (solid line) or less (dashed line) than 0.5, where event time equals zero at birth of the focal father' child. The figure shows clearly that log earnings are highly comparable in the pre-birth period, but diverge in the post-birth period.

To explore this observation more systematically using regression analysis, we first estimate equation (4), which conditions on individual characteristics as well as year-since-birth fixed effects, in a reduced-form regression that replaces C_i (competitor share taking the paternity leave) with S_i (competitor share eligible for it). We display the estimated coefficients in Figure 4, which clearly shows that the focal fathers with a higher share of leave-eligible competitor births are on a better earnings trajectory than their lower share counterparts but only in the periods following a child's birth. The positive earnings effect of having a high share of leave-eligible competitors is increasing until event year 4, after which it stabilizes. That this effect is not short-lived—e.g. focal father's earnings jump at year 1 but then fall back down when the competitors return from paternity leave—suggests that it is not an artefact of doing more work to make up for the competitors' temporary absence. Rather, a focal father with a high share of leave-eligible competitors seems to be put on a different earnings path than another focal father with a low share of leave-eligible competitors.

To facilitate subsequent discussion, we now estimate equation (5) so that the coefficient β reflects the average effect for up to 7 years after the child's birth. The OLS estimates indicate higher post-birth earnings for focal fathers when a larger share of their competitors take the paternity leave (Table 2, Panel A, columns 1 and 2). One concern may be that the competitors' leave taking is endogenous to the focal father or group characteristics. For instance, in a given contest, stronger contestants (i.e. those who are already winning), knowing that they will be ahead anyway, may be more at ease in taking leave than weaker ones. Moreover, leave-taking may be more common in contests where leave is unlikely to matter much in determining workers' ranking (wages) than in contests where leave is likely to matter. In both scenarios above, the OLS estimate will be biased downward.

To address this issue, we exploit variation in competitors' policy eligibility (driven by their child birthdates, which is plausibly exogenous to the focal father) and show the intention-to-treat (ITT) effect of competitor eligibility on focal fathers. Estimates in columns 3 and 4 show that as the share of policy eligible competitor births increases by 0.37 (1 *SD* in the sample)—which is equal to an additional two out of five competitor births (the sample average)—a focal father makes 1.1 percent higher earnings annually on average in the 7 years following his child's birth than he would have done otherwise.

The corresponding IV estimates (Table 2, columns 5 and 6) demonstrate the local average treatment effect (LATE) of being exposed to a larger share of competitor leave than otherwise because of the 1993 policy.[27] Column 6 suggests that as the share of leave-taking competitors increases by 0.32 (1 *SD* in the sample), the post-child earnings of a focal father are on average 2.4 percent higher than otherwise. The fact that the IV estimates here are larger than the OLS estimates (columns 1 and 2) suggests prevalence of endogenous leave taking such as the two scenarios discussed above.

Like in any IV analysis, it is important to interpret our IV estimates as the LATE, in this context the causal effect of competitors' leave (C_i) on focal fathers in the marginal contests where C_i is shifted because of our instrument (i.e. competitors' policy eligibility, S_i). To understand whether and along what dimensions the marginal contests may differ from the average contests, Table A1 compares the observable characteristics of "compliers" with that of average focal fathers in our sample, showing that compliers are largely similar to the overall sample except along a few dimensions.[28] For instance, compliers are more likely to be in a smaller competition group and work in the public sector compared to the average focal father in the sample. These characteristics of compliers should thus be taken into account when interpreting our IV estimates. In contrast, the

[27] The first stage corresponding to columns 5 and 6 shows a coefficient (SE) of 0.370 (0.015).

[28] "Compliers" are defined as focal fathers in contests in which "high S" (defined as $S \geq 0.5$ (median in sample)) results in "high C" (defined as $C \geq 0.32$ (median in the sample with $S \geq 0.5$)). Compliers' mean observable characteristics can be estimated by identifying the shares of always-takers and never-takers in the sample and using those shares to weight the observable characteristics of always-takers and compliers combined and that of always-takers alone, respectively (see, e.g. Almond and Doyle (2011) for a more complete description).

ITT effect speaks to the overall sample, and hence our discussion will focus on it for the remainder of the paper.

By choice of our observation window (symmetric around the policy cutoff date), around half the focal fathers have a child before (after) the cutoff and are thus leave ineligible (eligible). Because our empirical strategy relies on shifting the focal fathers' competitors in and out of leave eligibility while fixing the focal fathers' own eligibility statuses, the competition effect can be identified for either leave-ineligible or eligible focal fathers (although the magnitude need not be symmetric). To check this, we estimate a modified version of equation (5) where we interact our main regressors with focal father's own leave eligibility status. Results are reported in panel B of Table 2. As shown, the competition effect is invariant to the focal father's own leave eligibility status.

4.2 Robustness, placebo and heterogeneity analysis

As mentioned in Section 3.1, our main sample is restricted to focal fathers aged 25 to 36 (*P10* to *P90* of the age distribution of new fathers) to keep the sample more homogenous. However, our main estimates (Table 2) can be replicated when we do not impose such age restriction. As shown in Table A2, the competition effects remain very similar to that shown in Table 2.

In Table 3, we conduct further robustness checks for our primary competition effect. Column 1 replicates our main reduced-form estimate (column 4 in Table 2). In column 2, we add a birth year-month specific linear trend in event time (i.e. 18 different trends in year-since-birth), to allow for the possibility that focal fathers having a child at different points in time are on different earnings paths. This inclusion barely affects our results.

Next, in column 3, because the number of births within the contest itself may affect firm performance, we allow for linear trends in event time that differ by the number of competitor births in the group (1, 2, 3, 4, and 5+), a control to whose inclusion our estimate is invariant. Then, to address any concerns about the strategic timing of births, e.g. through changes in the timing of induction and cesarean section procedures (see Gans and Leigh 2009), we exclude from our sample the contests in which any birth (to focal father or competitor) occurs in March or April of 1993 (column 4).[29] The competition effect is still present and similar to that identified for our main sample.

We next perform a placebo analysis that investigates whether the share of leave-eligible competitors for a hypothetical reform enacted on April 1, 1992 (exactly one year before the actual reform date) would have similarly impacted the focal fathers.[30] As Table 4 shows, the share of leave-eligible competitors for this placebo reform has no noticeable impact (columns 1 and 2) regardless of the focal father's own eligibility status for it (columns 3-4 and 5-6). This confirms that our main findings, reported in Table 2, do indeed result from the 1993 paid paternity leave policy and not from a higher share of competitors having a child in or after April.

So far, we focused on the average effect across all firms. However, the competition effect (operating via paternity leave) is unlikely to be present in every workplace. For instance, in places where wage increases are mechanical and determined based on seniority and education there is little scope for the competition effect to operate. To explore this possibility, we test whether the estimated competition effects are larger among contests and competition groups in which the

[29] Since the reform was announced within 6 months prior to its implementation, there is no scope for strategic timing of *conception* (as opposed to birth) here. Also Figure 1 displays no notable changes in the frequency of births after April 1993.
[30] The sample of focal fathers for this exercise is constructed similarly to that for our main analysis; that is, male workers having a parity one child born within 9 months on either side of the (placebo) reform date of April 1, 1992.

dispersion of expected prizes—that is, the spread between W_H and W_L in Section 2.2—is largest, as suggested by Lazear and Rosen (1981). To proxy the expected prize relevant for each contest, we use the earnings growth from year -2 to 0 (since the focal father's child birth) of each male worker (aged 25 to 36) employed in the focal father's firm. We then classify the focal fathers into low vs. high dispersion groups based on the spread of the male employees' earnings growth in their firms. In Table 5, we consider two kinds of dispersion measures, *SD* (columns 1 and 2) and *P90* to *P50* (columns 3 and 4), and find that with either measure, the estimated competition effects are larger for focal fathers working in firms where the earnings growth of workers is more dispersed and the competition aspect is likely to be more relevant. Indeed, Figure 5 shows that firms with high dispersion of workers' earnings growth disproportionately come from the business and finance sector, consistent with the characterization of that sector in Bertrand et al. (2010) and Goldin (2014).

4.3 Total and direct effects

We next investigate the total effect of the focal father's own leave on his earnings. As mentioned above, as long as the competition channel remains open, research designs that exploit an exogenous shift in the focal father's own leave necessarily identify a combination of the direct and competition effects. This is because a shift in the focal father's own leave—even if exogenously induced by the policy—activates both a direct effect (absolute skill loss) and a competition (rank) effect. In contrast, an exogenous change in competitor leave creates only a competition effect. Below, we first identify the total effect, by estimating the effect of the focal father's own leave on his earnings. After that, we isolate the direct effect, by estimating the effect of the focal father's own leave, while shutting down the competition channel (by conditioning on the relative leave eligibility status of the focal father and his competitors).

To derive the total effect, we rely on the conditional randomness of the focal fathers' children's birth dates just around the policy cutoff date. To do so, we employ the following equation:

$$(6)\quad y_{it} = \phi_i + \psi_{\tau(it)} + \gamma Post_{\tau(it)} \times Z_i + \delta Post_{\tau(it)} \times f(X_i) + \boldsymbol{G}_{it}\lambda + e_{it},$$

where Z_i indicates focal father i's own eligibility for the paid paternity leave (i.e. child's birth date being after the policy cutoff date). Our focus is on γ, which designates the total effect of the focal father's leave eligibility on post-child earnings conditional on individual fixed effects (ϕ_i) and year-since-birth fixed effects ($\psi_{\tau(it)}$). The variable X_i denotes the expected birth week for focal father i's child, while $f(X_i)$ is a vector that flexibly accounts for the effects of X_i on earnings growth.[31] We also include a vector $\boldsymbol{G}_{it}$ that incorporates time-variant individual characteristics such as a cubic polynomial in age. We estimate equation (6) based on focal fathers with the child's expected birth week (X_i) within alternative windows on either side of the policy cutoff date.

As shown in panel A, Table 6, the total effect of the focal father's own leave eligibility (corresponding to "γ" in equation (6)) is around -2.8 to -3.3 percent. This total effect is similar in magnitude to that reported in previous studies of the Norwegian paternity leave reform. For instance, Rege and Solli (2013) report ITT estimates on earnings of -1 to -3 percent based on a DID design whereas Dahl et al. (2014) show a statistically insignificant effect size of -1.8 percent in an RDD approach.[32]

[31] In our baseline, we specify $f(X_i)$ as linear trends in X_i that are allowed to differ to the left (before) and right (after) of the policy cutoff date. For robustness, we also consider linear and quadratic trends in X_i.

[32] As mentioned above, Dahl et al. (2014) focus on relatively small firms where only one male worker has a child birth within 6 months on either side of the policy cutoff date, which may explain the smaller total effect of own leave, compared to Rege and Solli (2013)'s or our estimates.

Next, to disentangle the direct effect from the total effect, we re-estimate equation (6) while conditioning on $Rank_i$ defined as the difference between Z_i (the focal father's own leave eligibility) and S_i (the share of leave eligible competitors).[33] As illustrated in Figure 6, this $Rank_i$ variable is centered at zero and distributed between -1 (focal father ranked highest) and 1 (focal father ranked lowest), where 0 indicates symmetric eligibility for the focal father and his competitors. As panel B, Table 6 shows, once the competition (rank) channel is shut down, the effect of Z_i converges to zero, implying little direct effect associated with own leave. In Table A3, we demonstrate the robustness of these results to using linear (Panel A) and quadratic (Panel B) trends of expected birth week in place of the $f(X_i)$ used in our baseline.

Taken together, the above results confirm prior research evidence based on the 1993 reform that the total effect of the new father's own leave on his earnings is negative. More important, however, our results provide the novel insight that, in our context at least, the total effect on earnings is driven largely by the competition (rank) effect rather than the direct effect of own leave.[34]

5. Discussion

We focus on male-to-male comparison—i.e. between fathers having a child within the same pre-specified window—to stress the mechanics of the competition effect. However, the logic of the competition effect is more general. In particular, it suggests that if different leave statuses of workers are for reasons orthogonal to their abilities, then policies that remove the between-worker

[33] More precisely, we condition on $Post_{\tau(it)} \times Rank_i$. The level effect of $Rank_i$ is subsumed in the individual fixed effect ϕ_i.

[34] The absence of direct effect here may not be extrapolated to other situations where workers are on leave for substantially longer duration as in the case of maternity leave in Austria and Germany, for instance.

asymmetry in leave positions should eliminate the earnings penalty attached to leave, at least the part that operates via the competition channel. These may include universal paternity leave or normalization of remote working.

In the context of the 1993 policy of paid paternity leave quota, one question that naturally arises is whether the policy-induced paternity leave taken by new fathers can impact on other co-workers in the firm, and not just their competitors (i.e. comparable male workers also fathering a child during the same observation window). To investigate this, for each firm, we identify all male workers aged 25 to 36 with a child born during the 18-month window around April 1, 1993 ("fathers") and compute the share of births occurring after the policy cutoff date. We then examine—in a specification similar to equation (5)—how this share affects the earnings of different groups of male and female co-workers overlapping in the same firm with the fathers above during 1990-1992 (pre-period). To reduce the influence of noise from worker turnover, we restrict our attention to co-workers in the prime working age (25-49 as of 1992) with earnings above one social security basic unit during 1990-1992. We then examine how the share of leave-eligible fathers constructed above affects the earnings development of their male and female co-workers from the pre (1990-1992) to the post (1993-1996) reform periods.

In panel A of Table 7, we first look at firms with just one birth to a male worker (aged 25 to 36). Column (2) shows that the male worker's paternity leave eligibility has insignificant positive effect on "mothers", defined as female workers having a childbirth during the same 18-month window. For male (female) co-workers who are non-fathers (non-mothers), there is hardly any impact (columns (3) and (4)). Although the variation we exploit here comes from a male employee's paternity leave, the lack of significant general spillover effects echoes the findings of

Gallen (2018) and Brenøe et al. (2020) who examine maternity leave's spillover effects on co-workers using different reforms in Denmark.

In panel B, we then focus on firms with at least 2 births (to male workers aged 25 to 36), similar to our main analysis. In column (1), we find that post-1993 earnings of male workers with a child birth during the 18-month window are higher if they are exposed to a higher share of leave-eligible fathers.[35] Although the definition of fathers here is broader than what we used for our main analysis, the effect on this group of co-workers is in line with the competition effect we established earlier (Table 2). In contrast, we find little effect on co-workers who are mothers, non-fathers, or non-mothers (columns (2)-(4)).

Overall this analysis shows that the competition effect established above is driven by the relative comparison of the contestants (i.e. comparable male workers fathering a child around the reform implementation), rather than a generalized spillover effect affecting all workers in the firm. That a given father benefits from the leave of his competitors resembles the positive earnings effect caused by the unexpected death of a co-worker on the incumbent workers in the same occupation, as investigated in Jäger and Heining (2019). However, the earnings boost for co-workers in Jäger and Heining (2019) arises from the firm's inability to perfectly substitute the deceased worker—who permanently exits the firm—and the resulting increase in the value of the remaining co-workers to the firm, whereas in our setting, the competition effect concerns the ranking or relative standing of the incumbents—who all return to the firm once the leave period has ended. Therefore, the drivers of the earnings effect in the two studies are very different.

[35] For column (1), we construct the share S among all other fathers, leaving out the worker at hand.

6. Conclusions

Using variation induced by Norway's 1993 paid paternity leave quota together with the child birth date composition within pre-specified competition groups, we illustrate the competition effect—i.e. relative leave status of workers affecting their relative standing inside the firm—as an important yet unexplored driver of the career costs associated with paid leave. The role of the competition channel has largely been overlooked thus far as the literature on career interruptions has mainly concentrated on human capital and absolute productivity as a driver of worker earnings. In this paper we make the novel point that any credible research design that exploits exogenous variation in comparable workers' leave status and focuses on the effect of a worker's *own* leave on his own earnings is likely to identify a combination of the direct (skill depreciation) and competition (rank) effects. Indeed, we find that the total effect of own leave is very similar in magnitude to the competition effect, suggesting that the direct earnings effect resulting from own leave is negligible, at least in this context.

Although we focus on male-to-male comparison to facilitate clean identification of the competition effect, the logic of this effect should have broader implications. The European Council has recently adopted a *Directive on work-life balance for parents and carers* (EU Council Directive 2019/1158), mandating 2 months' paid and non-transferable parental leave for the father. Our analysis suggests that while such policies—to the extent that they normalize leave taking among all fathers—may be effective in eliminating the earnings penalty for leave-taking men, as long as the length of leave (conditional on child birth) remains skewed towards women than men, the playing field is unlikely to be levelled for female and male workers competing in promotions and career advancement. Therefore, the duration of paternity leave needed for equal progression of women and men on the firm hierarchy will be a question worthwhile investigating in future research. In addition, understanding how a radical expansion of remote work—in the aftermath of Covid-19—may change the earnings impact of leave-taking will be highly interesting.

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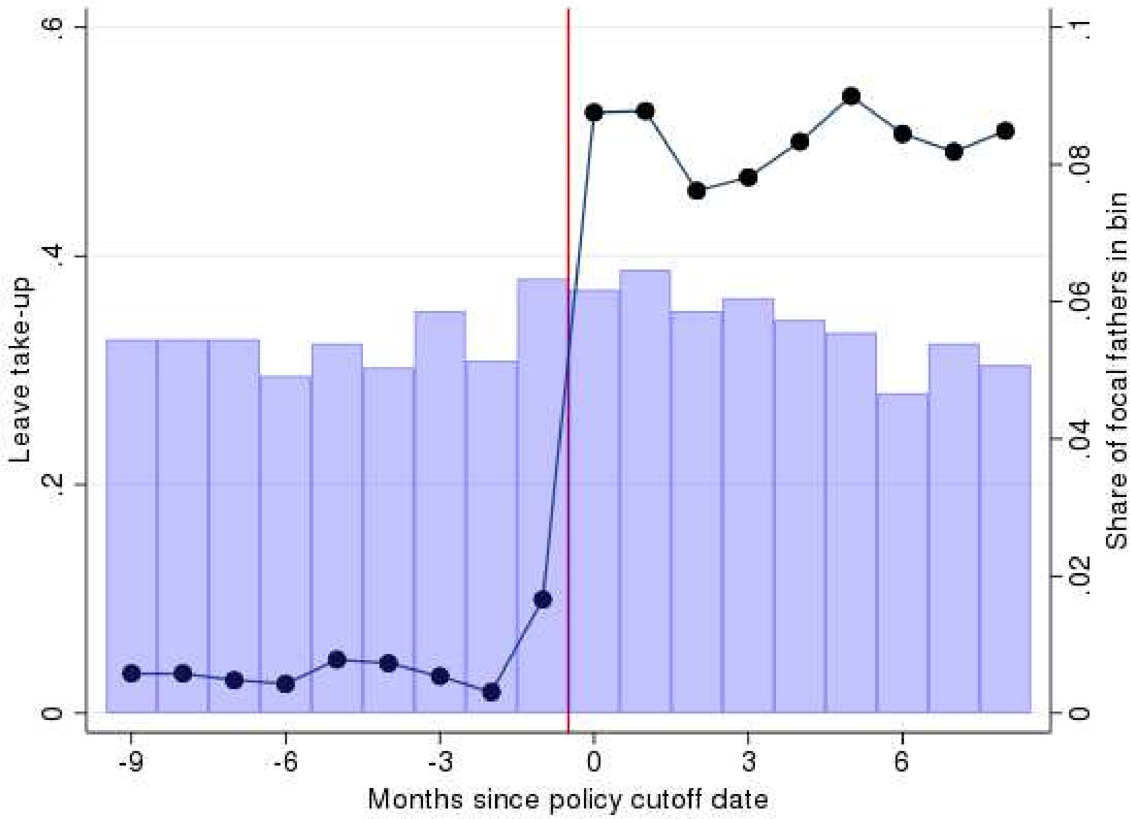
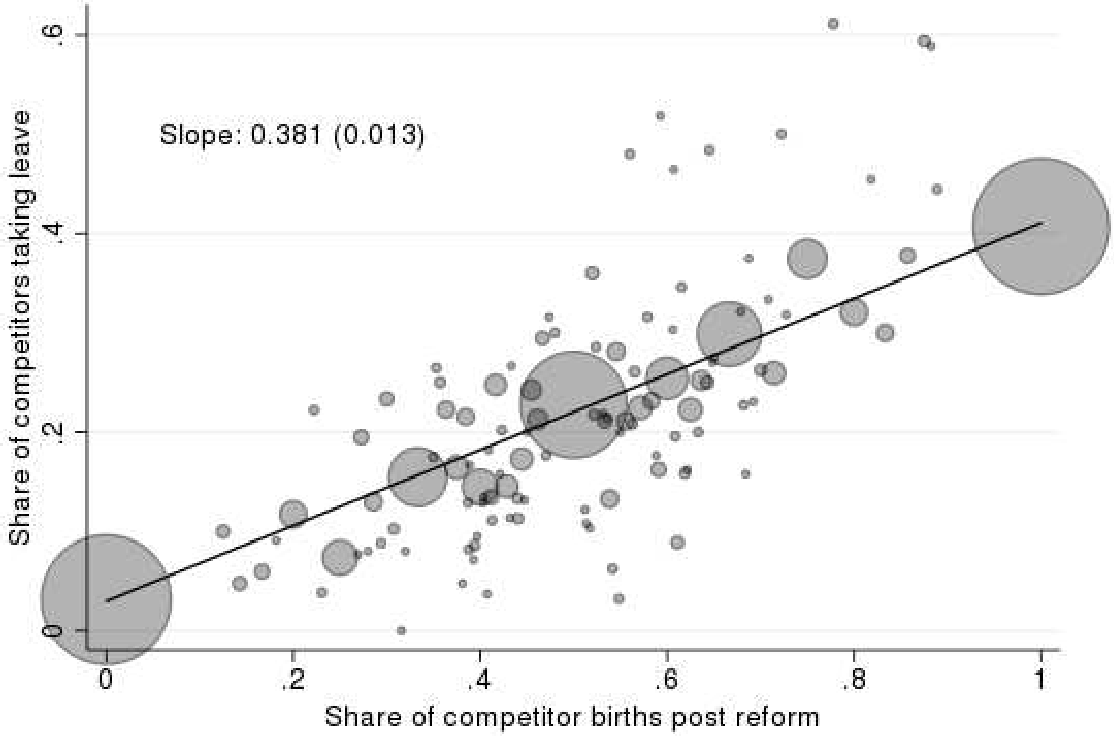
Figure 1. The new paid paternity leave quota and parental leave among focal fathers



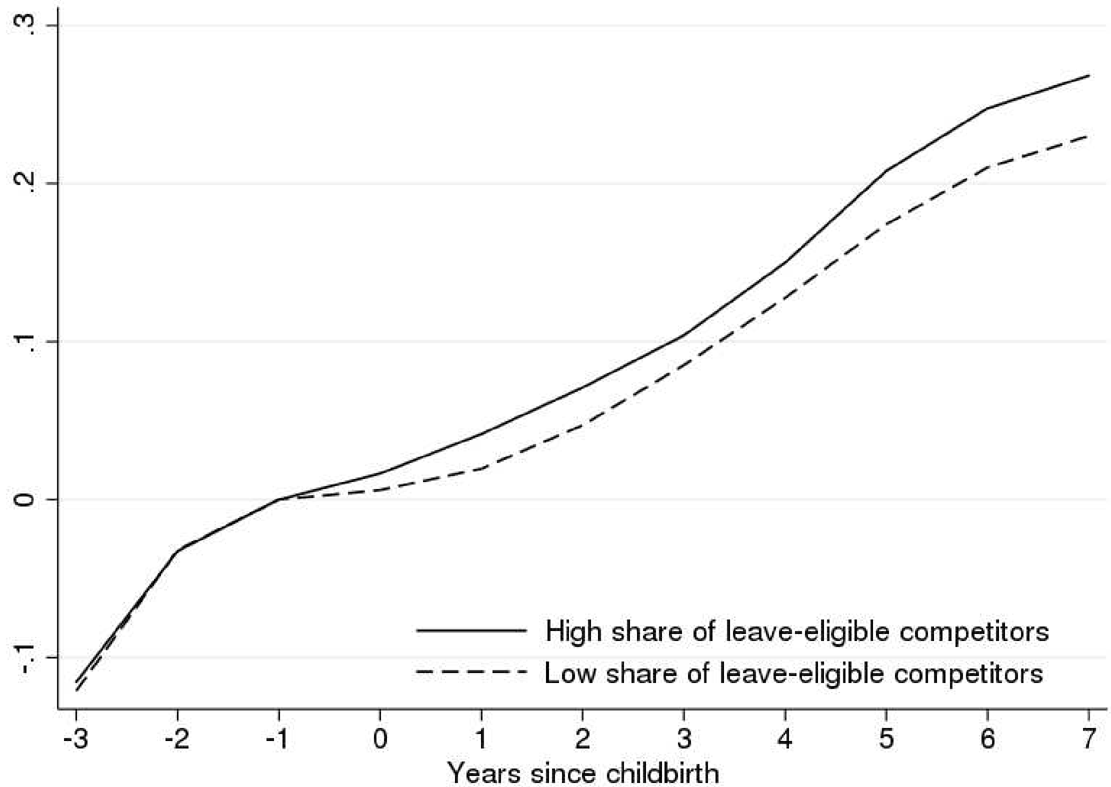
Notes: The connected dots in the figure plots the share of leave-taking focal fathers (measured on the left axis) by calendar month relative to the implementation month of the new paternity leave policy (April 1993). Focal fathers are defined as men who fathered their first child in the 18 month window around the introduction of the paid paternity leave quota. The bars in the figure represent the distribution of focal fathers in our sample (measured on the right axis) by child birth month. *Data source*: Norwegian birth register and social security records.

Figure 2. Share of competitors taking paternity leave vs. share of competitors eligible for the new paid paternity leave quota



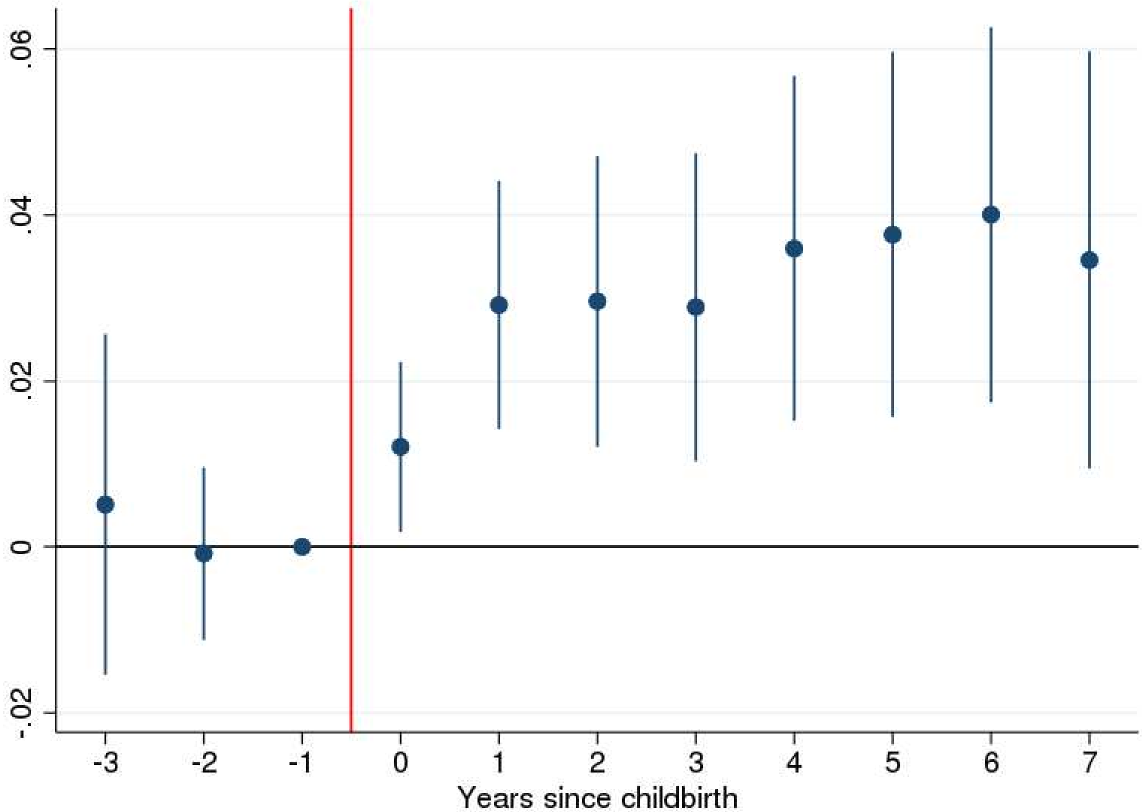
Notes: The figure plots the share of focal fathers' competitors taking parental leave against the share of competitors eligible for the paid paternity leave quota. The size of each circle is proportional to the number of focal fathers with the given share of leave-eligible competitors. The figure also reports the slope of the best-fit (weighted) line along with the standard error of the estimate (in parentheses). The figure is based on our main sample of focal fathers and their competitors. Focal fathers are defined as men who fathered their first child in the 18 month window around the introduction of the paid paternity leave quota on April 1st, 1993. Competitors are co-workers who also fathered a child in the 18 month window and have similar age and education as the focal father. *Data source*: Norwegian birth register, social security records, and matched employer-employee data.

Figure 3. Raw trends in log earnings of focal fathers with high versus low share of leave-eligible competitors



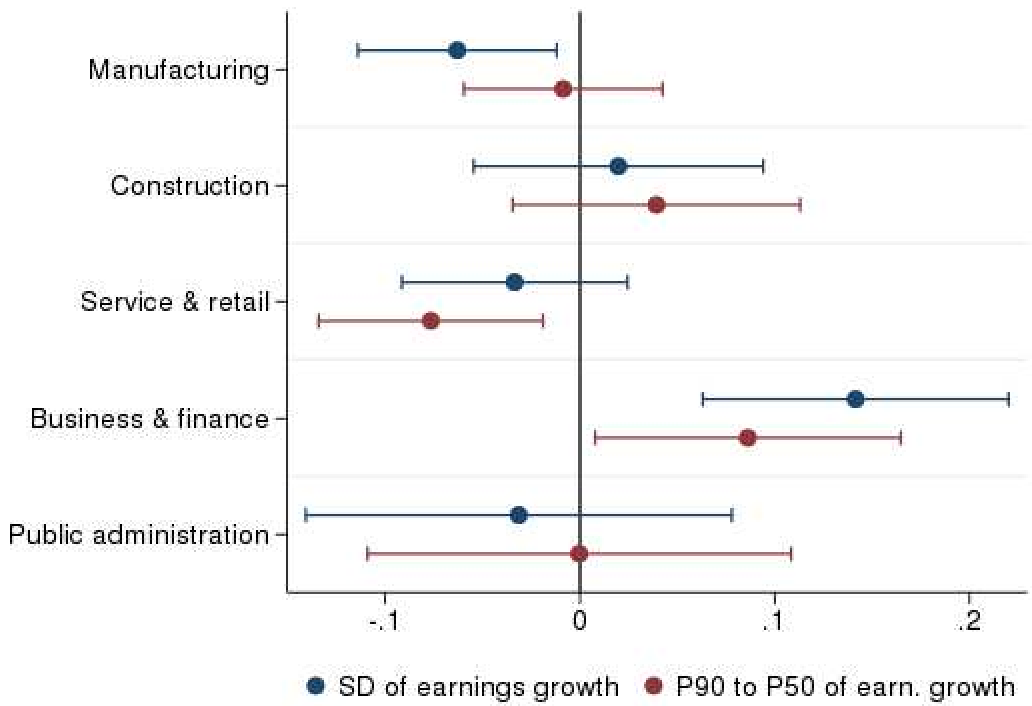
Notes: The figure presents average log earnings of focal fathers relative to the year before childbirth (i.e., year-since-birth = -1). The solid line represents focal fathers with a high share of leave-eligible competitors (i.e., share of leave-eligible competitors being at least 0.5). The dashed line represents focal fathers with a low share of leave-eligible competitors (i.e., share of leave-eligible competitors being below 0.5). Focal fathers are defined as men who fathered their first child in the 18 month window around the introduction of the paid paternity leave quota on April 1st, 1993. Competitors are co-workers who also fathered a child in the 18 month window and have similar age and education as the focal father. *Data source*: Norwegian birth register, tax records, and matched employer-employee data.

Figure 4. Effect of competitor leave eligibility on focal fathers' log earnings: Event study



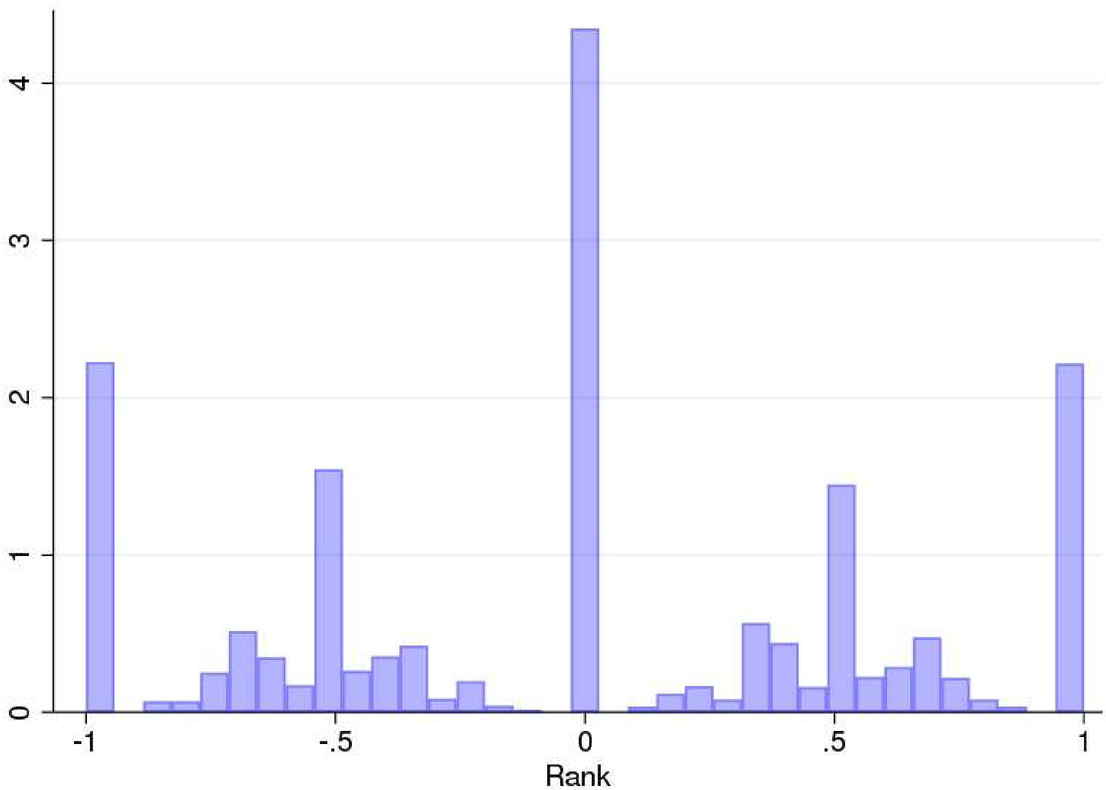
Notes: This figure displays the event year specific effects of the share of leave-eligible competitors on log earnings of focal fathers, estimated from equation (1). All effects are relative to the year before childbirth (i.e., year-since-birth = -1). The vertical lines indicate 90 percent CIs. The regression includes individual FE, event time FEs and a cubic polynomial in age. Focal fathers are defined as men who fathered their first child in the 18 month window around the introduction of the paid paternity leave quota on April 1st, 1993. Competitors are co-workers who also fathered a child in the 18 month window and have similar age and education as the focal father. *Data source*: Norwegian birth register, tax records, and matched employer-employee data.

Figure 5. Industry characteristics of firms with high vs. low dispersion of workers' earnings growth



Notes: This figure examines the industry characteristic of focal fathers' firms with high vs. low dispersion of worker earnings growth. Focal fathers are defined as men who fathered their first child in the 18 month window around the introduction of the paid paternity leave quota on April 1st, 1993. The figure plots the coefficients from univariate regressions of a high dispersion indicator on separate industry dummies. The horizontal lines indicate 95 percent CIs. We classify firms into the high (above median in the sample) dispersion group based on the spread of the earnings growth of male workers aged 25 to 36 from year -2 to 0 (since the focal father's child birth). We consider two dispersion measures, *SD* and *P90* to *P50*. The included industry indicators represent the main industries of the firms in the main sample. *Data source*: Norwegian birth register, tax records, and matched employer-employee data.

Figure 6. Distribution of focal fathers' rank, based on focal fathers' and competitors' eligibility for the paid paternity leave quota



Notes: This figure shows the distribution of focal father's *Rank* within the contest. We define *Rank* as *Z* – *S*, where *Z* is focal father's own eligibility for the paid paternity leave quota and *S* is the share of focal father's competitors eligible for the policy. Focal fathers are defined as men who fathered their first child in the 18 month window around the introduction of the paid paternity leave quota on April 1st, 1993. Competitors are co-workers who also fathered a child in the 18 month window and have similar age and education as the focal father. *Rank* is centered at zero and distributed between -1 (focal father ranked highest) and 1 (focal father ranked lowest), where 0 indicates symmetric eligibility for the focal father and his competitors. *Data source*: Norwegian birth register, social security records, and matched employer-employee data.

Table 1. Characteristics of focal fathers

	Overall		By share of leave-eligible competitors < 0.5		By share of leave-eligible competitors ≥ 0.5	
	Mean	SD	Mean	SD	Mean	SD
Focal father						
Birth year of child	1992.68	0.47	1992.68	0.47	1992.68	0.47
Male child	0.53	0.50	0.52	0.50	0.53	0.50
Take-up of parental leave	0.27	0.44	0.27	0.44	0.27	0.44
Leave duration (days)	39.19	44.90	40.31	48.11	38.43	42.61
Age at birth of child	30.00	4.23	30.12	4.37	29.91	4.13
Earnings (in 1,000 NOK)	247.70	98.20	245.93	87.09	248.92	105.17
Education						
≤ 2 years of high school	0.23	0.42	0.24	0.43	0.22	0.41
High school diploma	0.67	0.47	0.66	0.47	0.67	0.47
≤ 4 year college degree	0.05	0.22	0.05	0.21	0.05	0.23
> 4 years of college	0.05	0.23	0.05	0.22	0.06	0.23
Focal father's competitors						
# competitors	4.63	6.91	4.88	8.06	4.46	5.98
# leave-eligible / # competitors	0.51	0.37	0.13	0.17	0.77	0.22
# leave-takers / # competitors	0.22	0.32	0.08	0.16	0.32	0.35
Focal father's firm						
Firm size	463.82	755.48	435.76	680.72	483.15	802.51
Public sector	0.15	0.35	0.13	0.34	0.16	0.36
Mean earnings in firm	238.57	70.70	238.20	58.52	238.82	78.00
Mean age in firm	39.21	3.91	39.22	3.85	39.20	3.96
Observations	3822		1559		2263	

Notes: The table presents summary statistics for our main sample of focal fathers, both overall and by low vs. high share of leave-eligible competitors. Focal fathers are defined as men who fathered their first child in the 18 month window around the introduction of the paid paternity leave quota on April 1st, 1993. Competitors are co-workers who also fathered a child in the 18 month window and have similar age and education as the focal father. Eligibility for the paid paternity leave quota is based on having a child after the introduction of the policy on April 1st, 1993. All variables are measured at the birth year of the focal father's child. *Data source*: Data on father's age at the birth of his child plus birth year and gender of the child comes from the Norwegian birth register. Data on take-up and duration of parental leave comes from social security registers. Data on earnings comes from tax registers. Earnings are deflated to 1993 Norwegian kroner using the CPI. Education data comes from national education registers. Firm data comes from employer-employee registers which we also use to link focal fathers to their competitors and to calculate average characteristics of competitors.

Table 2. Effect of competitor leave eligibility on focal fathers' log earnings

	Dependent variable: Focal father's log earnings					
	OLS		Reduced-form		IV	
	(1)	(2)	(3)	(4)	(5)	(6)
A. Overall effect						
$C \times Post$	0.0312**	0.0339***			0.0860***	0.0774***
	(0.012)	(0.012)			(0.027)	(0.026)
$S \times Post$			0.0327***	0.0295***		
			(0.010)	(0.010)		
B. Effect by focal father's own leave eligibility						
$C \times Post$	0.0456**	0.0477***			0.0922***	0.0826***
	(0.018)	(0.017)			(0.032)	(0.032)
$C \times Post \times I(Z=0)$	-0.0309	-0.0295			-0.0131	-0.0109
	(0.021)	(0.021)			(0.030)	(0.030)
$S \times Post$			0.0361***	0.0324**		
			(0.013)	(0.013)		
$S \times Post \times I(Z=0)$			-0.0070	-0.0059		
			(0.013)	(0.012)		
Cubic in age	no	yes	no	yes	no	yes
Observations	35871		35871		35871	

Notes: This table reports estimates of equation (5) based on our main sample of focal fathers, covering year-since-birth (YSB) from -3 to 7. Focal fathers are defined as men who fathered their first child in the 18 month window around the introduction of the paid paternity leave quota on April 1st, 1993. Competitors are co-workers who also fathered a child in the 18 month window and have similar age and education as the focal father. *C* denotes the share of focal fathers' competitors who are taking paternity leave. *S* denotes the share of focal father's competitors who are eligible for the paid paternity leave quota. *Post* is 1 for YSB of 0 to 7, and 0 for YSB of -3 to -1. Panel A shows the overall effect. Panel B shows the effect by focal father's own eligibility status for the paid paternity leave quota, where $I(Z=0)$ is an indicator taking the value of unity if the focal father is leave ineligible. Columns 1 and 2 report estimates from OLS regressions. Columns 3 and 4 report reduced-form estimates of equation (5). Columns 5 and 6 report IV estimates of equation (5), where the share of leave-taking competitors is instrumented by the share of leave-eligible competitors. The first stage corresponding to columns 5 and 6 shows a coefficient (SE) of 0.370 (0.015). All regressions include individual FEs and year-since-birth (event time) FEs. Even-numbered columns also include a cubic polynomial in focal father's age. Standard errors are clustered at the firm level. *p<0.10, **p<0.05, ***p<0.01. *Data source*: Norwegian birth register, tax records, social security records, and matched employer-employee data.

Table 3. Robustness of the competition effect

Dependent variable: Focal father's log earnings				
	Specification			
	(1)	(2)	(3)	(4)
$S \times Post$	0.0295***	0.0289***	0.0293***	0.0327***
	(0.010)	(0.010)	(0.010)	(0.012)
Event time (year-since-birth) FE	yes	yes	yes	yes
Cubic polynomial in age	yes	yes	yes	yes
Individual FE	yes	yes	yes	yes
Linear trend by birth year-month	no	yes	no	no
Linear trend by # of competitor births	no	no	yes	no
Drop March and April 1993	no	no	no	yes
Observations	35871	35871	35871	20768

Notes: This table reports reduced-form estimates of equation (5) based on our main sample of focal fathers, covering year-since-birth (YSB) from -3 to 7. Focal fathers are defined as men who fathered their first child in the 18 month window around the introduction of the paid paternity leave quota on April 1st, 1993. Competitors are co-workers who also fathered a child in the 18 month window and have similar age and education as the focal father. *S* denotes the share of focal father's competitors who are eligible for the paid paternity leave quota. *Post* is 1 for YSB of 0 to 7, and 0 for YSB of -3 to -1. Column 1 replicates our main specification (column 4 of Table 2). Column 2 adds birth year-month specific linear trends in event time. Column 3 includes number of births among competitors (1, 2, 3, 4, 5+) specific linear trend in event time (YSB). Column 4 is a donut specification dropping contests that include any births (to focal father or to competitor) occurring in March or April 1993. Standard errors are clustered at the firm level. *p<0.10, **p<0.05, ***p<0.01. *Data source*: Norwegian birth register, tax records, and matched employer-employee data.

Table 4. Competition effect from a placebo reform

Dependent variable: Focal father's log earnings						
	Overall sample		*"Ineligible" focal fathers*		*"Eligible" focal fathers*	
	(1)	(2)	(3)	(4)	(5)	(6)
$S \times Post$	-0.0000	-0.0029	0.0069	0.0020	-0.0076	-0.0080
	(0.009)	(0.009)	(0.013)	(0.013)	(0.013)	(0.013)
Cubic in age	no	yes	no	yes	no	yes
Observations	39721		19921		19800	

Notes: This table reports reduced-form estimates of equation (5), in which leave-eligibility is based on a placebo reform date of April 1st, 1992 (as opposed to the actual reform date of April 1st, 1993 in our main analysis). For this analysis, focal fathers are defined as men who fathered their first child in the 18 month window around the introduction of the placebo reform on April 1st, 1992. Competitors are co-workers who also fathered a child in the 18 month window and have similar age and education as the focal father. *S* is the share of competitors with child birth on or after April 1, 1992. *Post* is 1 for (focal father's) year-since-birth (YSB) of 0 to 7, and 0 for YSB of -3 to -1. "Ineligible" focal fathers are defined as focal fathers with child birth before April 1, 1992. "Eligible" focal fathers are defined as focal fathers with child birth before on or after April 1, 1992. All regressions include individual FEs and year-since-birth (event time) FEs. Even-numbered columns also include a cubic polynomial in focal father's age. Standard errors are clustered at the firm level. *p<0.10, **p<0.05, ***p<0.01. *Data source*: Norwegian birth register, tax records, and matched employer-employee data.

Table 5. Competition effect by the dispersion of expected prizes in the contest

	Dependent variable: Focal father's log earnings			
	Dispersion: *SD* of earnings growth among male coworkers		Dispersion: *P90* to *P50* of earnings growth among male coworkers	
	Low dispersion (1)	High dispersion (2)	Low dispersion (1)	High dispersion (2)
$S \times Post$	0.0197*	0.0366**	0.0115	0.0503***
	(0.011)	(0.017)	(0.013)	(0.016)
Observations	18040	17776	18051	17765

Notes: This table reports reduced-form estimates of equation (5), by the dispersion of expected prizes in the contest, based on our main sample of focal fathers, covering year-since-birth (YSB) from -3 to 7. Focal fathers are defined as men who fathered their first child in the 18 month window around the introduction of the paid paternity leave quota on April 1st, 1993. Competitors are co-workers who also fathered a child in the 18 month window and have similar age and education as the focal father. *S* denotes the share of focal father's competitors who are eligible for the paid paternity leave quota. *Post* is 1 for YSB of 0 to 7, and 0 for YSB of -3 to -1. Expected prize are defined as the earnings growth from (focal father's) YSB -2 to 0 of male workers aged 25 to 36 who are employed in the same plant as the focal father while we measure dispersion as the spread of this pre-birth earnings growth within each plant. Low dispersion is defined as below median of the stated spread measure, and high dispersion is defined as above median of the stated spread measure. We control for individual FEs, year-since-birth (event time) FEs, and a cubic polynomial in focal father's age. Standard errors are clustered at the firm level. *p<0.10, **p<0.05, ***p<0.01. *Data source*: Norwegian birth register, tax records, and matched employer-employee data.

Table 6. Total and direct effects of a focal father's own leave eligibility on his log earnings

Dependent variable: Focal father's log earnings				
	Window size			
	2x9 months	2x8 months	2x7 months	2x6 months
	(1)	(2)	(3)	(4)
A. Total effect (unconditional on rank effect)				
$Z \times Post$	-0.0311**	-0.0285*	-0.0333**	-0.0319*
	(0.014)	(0.015)	(0.016)	(0.017)
B. Direct effect (conditional on rank effect)				
$Z \times Post$	0.0015	0.0052	0.0002	-0.0074
	(0.018)	(0.019)	(0.021)	(0.022)
$Rank \times Post$	-0.0335***	-0.0349***	-0.0341***	-0.0249*
	(0.011)	(0.011)	(0.012)	(0.013)
Observations	31394	28303	24893	21549

Notes: This table reports estimates of equation (6), examining the effect of focal father's own leave-eligibility on his log earnings. Focal fathers are defined as men who fathered their first child in a narrow window around the introduction of the paid paternity leave quota on April 1st, 1993. We vary this window in columns 1 to 4, ranging from 9 to 6 months on either side of the policy cutoff date. Competitors are co-workers who also fathered a child in the 18 month window and have similar age and education as the focal father. *Z* indicates whether the focal father is eligible for the paid paternity leave quota or not. *S* denotes the share of focal father's competitors who are eligible for the paid paternity leave quota. *Rank* is defined as *Z* – *S*. *Post* is 1 for (focal father's) year-since-birth (YSB) of 0 to 7, and 0 for YSB of -3 to -1. Panel A shows results from regressions without conditioning on the rank effect. Panel B shows results from regressions that condition on the rank effect. We control for individual FEs, year-since-birth (event time) FEs, a cubic polynomial in focal father's age, and linear trends in expected week of child birth that are allowed to differ on either side of the policy cutoff date. Standard errors are clustered at the firm level. *p<0.10, **p<0.05, ***p<0.01. *Data source*: Norwegian birth register, tax records, and matched employer-employee data.

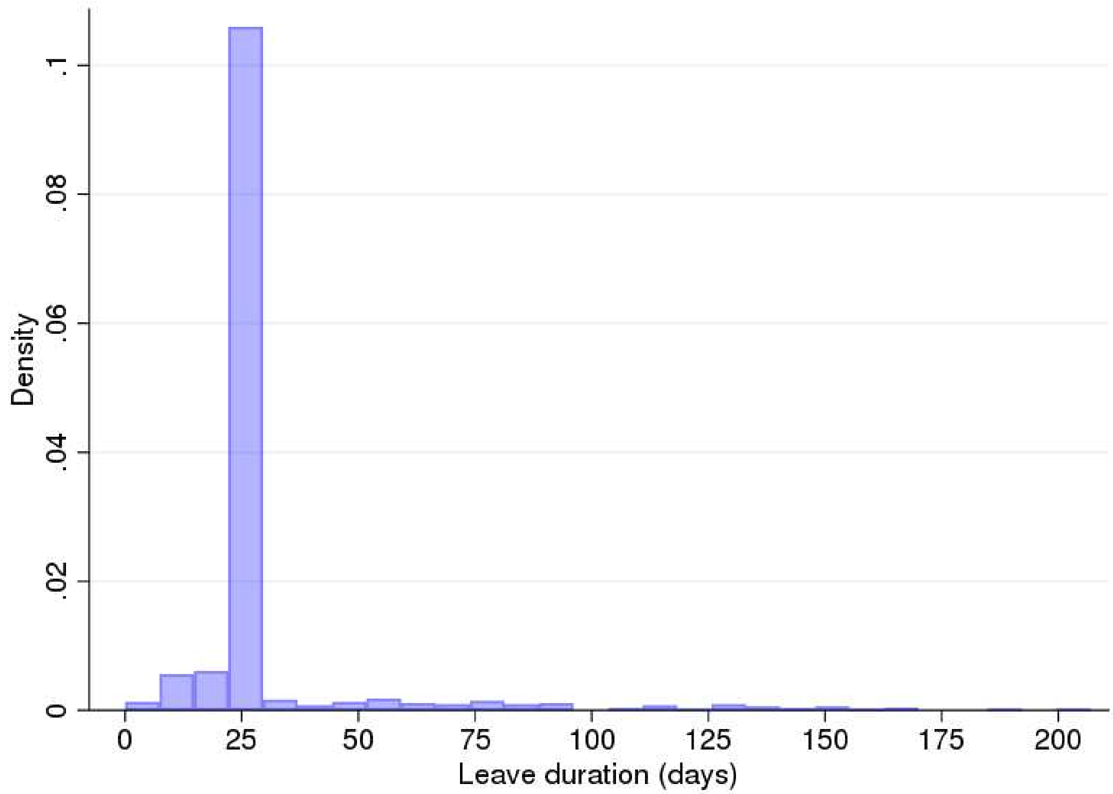
Table 7. Competition vs. spillover effect from leave eligibility of male co-workers in the firm

Dependent variable: Co-worker log earnings				
	Co-worker type			
	Fathers (1)	Mothers (2)	Non-fathers (3)	Non-mothers (4)
A. 1 birth case				
$S \times Post$	N/A	0.0235	0.0044	0.0040
		(0.018)	(0.004)	(0.004)
Observations		17523	276491	222470
B. 2+ birth case				
$S \times Post$	0.0122**	0.0038	0.0035	0.0032
	(0.006)	(0.015)	(0.004)	(0.006)
Observations	156414	69289	1208571	849111

Notes: This table reports estimates of from a regression of worker's own log earnings on the share (S) of leave-eligible male co-workers in the firm aged 25 to 36 and having a child in a 2x9 months window around the introduction of the paid paternity leave quota on April 1st, 1993. For fathers in column (1), S is defined as the leave-out mean of S. We follow individuals from 1990 to 1996, and *Post* is 1 for calendar years 1993 onwards. Panel A shows results for individuals working in firms in which only one male co-worker had a child in the specified window. Panel B shows results for individuals working in firms in which at least two male co-workers had a child in the specified window. We control for individual FEs, and a cubic polynomial in the individual's age. Standard errors are clustered at the firm level. *p<0.10, **p<0.05, ***p<0.01. *Data source*: Norwegian birth register, tax records, and matched employer-employee data.

Appendix A

Figure A1. Distribution of paternity leave duration for leave-taking focal fathers



Notes: This figure shows the distribution of paternity leave duration (in days) for leave-eligible focal fathers who take up leave. Focal fathers are defined as men who fathered their first child in the 18 month window around the introduction of the paid paternity leave quota on April 1st, 1993. For expositional reasons, the figure excludes observations of leave duration in the top 1 percentile. *Data source*: Norwegian birth register and social security records.

Table A1. Characteristics of compliers

	Complier Mean	Overall Mean
Focal father		
Birth year of child	1992.72	1992.68
Male child	0.55	0.53
Take-up of parental leave	0.30	0.27
Leave duration (days)	37.60	39.19
Age at birth of child	30.18	30.00
Earnings (in 1,000 NOK)	255.97	247.70
Education		
≤ 2 years of HS	0.18	0.23
HS diploma	0.66	0.67
≤ 4 yr college degree	0.09	0.05
>4 yrs of college	0.07	0.05
Focal father's competitors		
# competitors	3.21	4.63
# leave-eligible / # competitors (*S*)	0.94	0.51
# leave-takers / # competitors (*C*)	0.71	0.22
Focal father's firm		
Firm size	413.50	463.82
Public sector	0.18	0.15
Mean earnings in firm (in 1,000 NOK)	240.85	238.57
Mean age in firm	39.48	39.21

Notes: The table shows average characteristics for "compliers" and the overall sample of focal fathers. Focal fathers are defined as men who fathered their first child in the 18 month window around the introduction of the paid paternity leave quota on April 1st, 1993. Competitors are co-workers who also fathered a child in the 18 month window and have similar age and education as the focal father. "Compliers" are defined as focal fathers in contests in which "high S" (defined as $S \geq 0.5$ (median in sample)) results in "high C" (defined as $C \geq 0.32$ (median in the sample with $S \geq 0.5$)). Compliers' mean observable characteristics are estimated following the procedure described in Almond and Doyle (2011), where we first identify the shares of always-takers and never-takers in the sample and then use those shares to weight the observable characteristics of always-takers and compliers combined and that of always-takers alone, respectively. *Data source*: Data on father's age at the birth of his child plus birth year and gender of the child comes from the Norwegian birth register. Data on take-up and duration of parental leave comes from social security registers. All variables are measured at the birth year of the focal father's child. Data on earnings comes from tax registers. Earnings are deflated to 1993 Norwegian kroner using the CPI. Education data comes from national education registers. Firm data comes from employer-employee registers which we also use to link focal fathers to their competitors and competitors' characteristics.

Table A2. Effect of competitor leave eligibility on focal fathers' log earnings in sample of focal fathers without age restriction

	Dependent variable: Focal father's log earnings					
	OLS		Reduced-form		IV	
	(1)	(2)	(3)	(4)	(5)	(6)
A. Overall effect						
$C \times Post$	0.0296**	0.0354***			0.0805***	0.0648***
	(0.012)	(0.011)			(0.026)	(0.025)
$S \times Post$			0.0298***	0.0240***		
			(0.010)	(0.009)		
B. Effect by focal father's own leave eligibility						
$C \times Post$	0.0402**	0.0453***			0.0788**	0.0663**
	(0.017)	(0.016)			(0.031)	(0.029)
$C \times Post \times I(Z = 0)$	-0.0225	-0.0211			0.0036	-0.0031
	(0.020)	(0.019)			(0.029)	(0.028)
$S \times Post$			0.0299**	0.0253**		
			(0.012)	(0.012)		
$S \times Post \times I(Z = 0)$			-0.0003	-0.0027		
			(0.012)	(0.012)		
Cubic in age	no	yes	no	yes	no	yes
Observations	42042		42042		42042	

Notes: This table reports estimates of equation (5) based on a sample of focal fathers who are not restricted to ages 25 (p10) and 36 (p90), covering year-since-birth (YSB) from -3 to 7. Focal fathers are defined as men who fathered their first child in the 18 month window around the introduction of the paid paternity leave quota on April 1st, 1993. Competitors are co-workers who also fathered a child in the 18 month window and have similar age and education as the focal father. *C* denotes the share of focal fathers' competitors who are taking paternity leave. *S* denotes the share of focal father's competitors who are eligible for the paid paternity leave quota. *Post* is 1 for YSB of 0 to 7, and 0 for YSB of -3 to -1. Panel A shows the overall effect. Panel B shows the effect by focal father's own eligibility status for the paid paternity leave quota, where $I(Z = 0)$ is an indicator taking the value of unity if the focal father is leave ineligible. Columns 1 and 2 report estimates from OLS regressions. Columns 3 and 4 report reduced-form estimates of equation (5). Columns 5 and 6 report IV estimates of equation (5), where the share of leave-taking competitors is instrumented by the share of leave-eligible competitors. All regressions include individual FEs and year-since-birth (event time) FEs. Even-numbered columns also include a cubic polynomial in focal father's age. Standard errors are clustered at the firm level. *p<0.10, **p<0.05, ***p<0.01. *Data source*: Norwegian birth register, tax records, social security records, and matched employer-employee data.

Table A3. Total and direct effects of focal father's own leave eligibility on his log earnings: Robustness to alternative specifications

Dependent variable: Focal father's log earnings				
	Window size			
	2x9 months	2x8 months	2x7 months	2x6 months
	(1)	(2)	(3)	(4)
A. Linear controls of expected birth week				
A1. Total effect (unconditional on rank effect)				
$Z \times Post$	-0.0305**	-0.0283*	-0.0328**	-0.0305*
	(0.014)	(0.015)	(0.016)	(0.017)
A2. Direct effect (conditional on rank effect)				
$Z \times Post$	0.0022	0.0054	0.0008	-0.0060
	(0.018)	(0.019)	(0.020)	(0.022)
$Rank \times Post$	-0.0337***	-0.0349***	-0.0342***	-0.0250**
	(0.011)	(0.011)	(0.012)	(0.013)
B. Quadratic controls for expected birth week				
B1. Total effect (unconditional on rank effect)				
$Z \times Post$	-0.0308**	-0.0284*	-0.0329**	-0.0313*
	(0.014)	(0.015)	(0.016)	(0.017)
B2. Direct effect (conditional on rank effect)				
$Z \times Post$	0.0017	0.0054	0.0006	-0.0070
	(0.018)	(0.019)	(0.020)	(0.022)
$Rank \times Post$	-0.0334***	-0.0349***	-0.0341***	-0.0248*
	(0.011)	(0.011)	(0.012)	(0.013)
Observations	31394	28303	24893	21549

Notes: This table reports estimates of equation (6), examining the effect of focal father's own leave eligibility on his log earnings. Focal fathers are defined as men who fathered their first child in a narrow window around the introduction of the paid paternity leave quota on April 1st, 1993. We vary this window in columns 1 to 4, ranging from 9 to 6 months on either side of the policy cutoff date. *Z* indicates whether the focal father is eligible for the paid paternity leave quota or not. *S* denotes the share of focal father's competitors who are eligible for the paid paternity leave quota. Competitors are co-workers who also fathered a child in the 18 month window and have similar age and education as the focal father. *Rank* is defined as *Z – S*. *Post* is 1 for (focal father's) year-since-birth (YSB) of 0 to 7, and 0 for YSB of -3 to -1. Panels A1 and B1 show results from regressions without conditioning on the rank effect. Panels A2 and B2 show results from regressions that condition on the rank effect. All regressions include individual FEs, year-since-birth (event time) FEs, and a cubic polynomial in focal father's age. In Panels A1 and A2, we also include linear trends in expected week of child's birth. In Panels B1 and B2, we control for quadratic trends in expected week of child's birth. Standard errors are clustered at the firm level. *p<0.10, **p<0.05, ***p<0.01. *Data source*: Norwegian birth register, tax records, and matched employer-employee data.