The Impact of Paternity Leave on Long-term Father Involvement

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Abstract

Using Norwegian registry data we investigate how paternity leave affects fathers' long-term earnings. In 1993 Norway introduced a paternity quota of the paid parental leave. We estimate a difference-in-differences model which exploits differences in fathers' exposure to the paternity quota. Our analysis suggests that four weeks paternity leave during the child's first year decreases fathers' future earnings by 2.1 percent. Importantly, this effect persists up until our last point of observation when the child is five years old. The earnings effect is consistent with increased long-term father involvement, as fathers shift time and effort from market to home production. In an investigation of Norwegian time use data we find additional evidence for this hypothesis.

JEL-Code: D13, H31, J22.

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1. Introduction

During the last decades there has been an increased political and public concern over fathers' involvement in their children's lives. In the USA, for example, the Clinton Administration launched in 1995 a government-wide initiative to strengthen the role of fathers in families.[1] Moreover, American television watchers have frequently come across advertising campaigns encouraging male viewers to be more involved with their children.[2] This concern over father involvement has been fuelled by increasing empirical evidence suggesting that the involvement of a father in his children's lives is important for the children's cognitive and socio-emotional outcomes.[3]

In this paper we investigate whether paternity leave during the child's first year can increase long-term father involvement. We consider a father to be more involved with his child if he spends more time together with that child.[4] Paternity leave might affect a father's long-term involvement through at least two different mechanisms (Tanaka and Waldfogel 2007). First, a father caring for an infant child may facilitate father-child bonding. Second, paternity leave could make it easier for the father to be more involved as the child grows older by preventing the mother from gaining exclusive expertise in child caring during the child's first year.

There is a large and recent economic literature investigating impacts of maternity leave.[5] However, the empirical evidence on paternity leave is scant. Even if not conclusive, this study provides some of the first evidence that paternity leave has a causal effect on father long-term involvement.[6] This is important because it suggests that paternity leave policies have implications for child well-being (Han, Ruhm and Waldfogel 2009). The policy relevance of our findings is highlighted by a recent resolution in the European Parliament. In March 2010, the European Parliament adopted a directive stipulating the minimum

[1] Clinton, 1995: Memorandum for the Heads of Executive Departments and Agencies: Supporting the Role of Fathers in Families. White House, June 16, 1995.

[2] See for example advertisements from the National Father Initiative at http://www.fatherhood.org/media/PSAs/tv.asp.

[3] See for example Lamb (2010) and Tamis-Lemonda and Cabrera (2002).

[4] Perhaps unfairly, a father who never sees his children because he works long hours in order to make enough money for their college education is not characterized as an involved father.

[5] Se for example Baker and Milligan (2008a, 2008b, 2010), Carneiro, Løken and Salvanes (2009a, 2009b), Dustmann and Schönberg (2008), Ruhm (1998, 2004), Schönberg and Ludsteck (2007).

[6] There exist several studies documenting an association between paternity leave and father involvement. Using US data, Nepomnyaschy and Waldfogel (2007) demonstrate that longer paternity leave at the time of the childbirth is associated with more child care-taking activities nine months after the birth. Similar results are found in Tanaka and Waldfogel (2007) utilizing UK survey data and Haas and Hwang (2008) utilizing Swedish survey data. See recent review of empirical findings in Haas and Hwang (2008).

requirements for parental leave, including a non-transferrable paternity quota of four weeks.[7] Our study suggests that paternity leave has the expected positive effect on long-term father involvement and that implementing paternity quotas of the parental leave is an adequate policy action to support the role of fathers in families.

Estimating a causal relationship between paternity leave and father involvement faces two main challenges: concerns of omitted variable bias and the scarcity of appropriate data. Omitted variable bias arises if a father's decision to take paternity leave is correlated with unobservables that also affect father involvement, such as the father's preferences for spending time with his child. To circumvent the most obvious forms of omitted variable bias, our empirical strategy utilizes the Norwegian introduction of a paternity quota in 1993. From 1993, four weeks of the total of 42 weeks of paid parental leave were reserved exclusively for the father. With few exceptions the family would lose those four weeks of paid parental leave if these were not taken by the father. The introduction of the paternity quota led to a sharp increase in uptake rates. In our sample of full-time working fathers, the utilization rate was less than three percent prior to 1993, but had increased to about 60 percent already by 1995.

Investigating the relationship between paternity leave and father involvement is also constrained by data availability. The task requires appropriate indicators for father involvement throughout the child's life. In this study we obtain these indicators by utilizing a comprehensive, longitudinal registry database containing annual records of earnings for every person in Norway. If the paternity quota increased long-term father involvement, then we should expect a reduction in fathers' long-term earnings, as they shift time and effort from market to home production (Becker 1985).[8] We supplement our investigation of registry data with analyses of data from Norwegian time use surveys.

In our main empirical investigation we estimate the effects of the paternity quota on fathers' earnings. We estimate a difference-in-differences (DD) model which exploits differences in fathers' exposure to the paternity quota. Our analysis suggests that four weeks paternity leave during the child's first year decreases fathers' future earnings by 2.1 percent. This effect persists up until our last point of observation which is when the child is five years old.

The negative effect of paternity leave on long-term earnings is consistent with the idea that the father is spending less time at work and more time together with his child. However,

[7] European Union: Council Directive 2010/18/EU.

[8] Notably, reduced labor supply (and possibly lower productivity) has a *direct* negative effect on short and long-term earnings, in addition to an *indirect* negative effect on long-term earnings through reduced human capital accumulation.

there are also several other reasons why the quota could affect fathers' earnings. For example, absence from work while being on paternity leave reduces accumulation of work experience and work related human capital. Alternatively, taking time off from work to be on paternity leave may serve as a signal of being more family-oriented rather than career-oriented. Employers may consider such employees as being less devoted and reliable, thus reducing the likelihood of their giving promotions and pay raises.

Unfortunately, our registry data does not allow us to distinguish between the different mechanisms for how the paternity quota affects fathers' earnings. Consequently we turn to time diaries from the Norwegian Time Use Surveys in 1990 and 2000 in order to provide more direct evidence for the effect of the paternity quota on father involvement. Using a similar difference-in-differences approach, we find that fathers spent significantly less time working and more time together with their children after the paternity quota was implemented. Admittedly, an important limitation of this analysis is that we only observe fathers' time use in 1990 and 2000. This makes it impossible to closely connect the changes in time use to the introduction of the paternity quota. Nevertheless, *together* with our analysis of registry data, which provides convincing evidence of a causal effect of the paternity quota on earnings, the time use analyses suggest that the paternity quota had an impact on long-term father involvement.

The remainder of the paper is organized as follows: In Section 2 we give a brief overview of the paternity quota and other relevant family policies. Section 3 describes our registry data, and in Section 4 we discuss our empirical strategy. Section 5 presents our results. In Section 6 we investigate mechanisms using data from time use surveys. Section 7 concludes.

2. The Paternity Quota

On April 1st 1993, Norway introduced a paternity quota of the paid parental leave. Four weeks of the total of 42 weeks of paid parental leave were reserved exclusively for the father.[9] With few exceptions, the family would lose those four weeks of paid parental leave if not taken by the father. Apart from the four weeks paternity quota and three weeks prior to and six weeks after birth, which were reserved for the mother for medical reasons, parents

[9] Alternatively, parents could take 52 weeks of parental leave at 80 percent pay. Earnings above 6*"Basic Amount" (around €19 000 in 2010) are not compensated by the government. Around 17 (48) percent of all women (men) above 17 years of age earn more than this earnings ceiling. However, most employers (private and public) compensate earnings above this ceiling.

could share the parental leave between them as they desired.[10] While paid maternity leave was only contingent on the mother working at least 50 percent of full-time prior to birth, paid paternity leave was contingent on both parents working at least 50 percent. Income compensation was based on the earnings of the person on leave, but fathers' income compensation was reduced proportionally if the mother did not work full-time prior to birth.[11]

The introduction of the paternity quota led to a sharp increase in uptake rates. Based on our analytical sample of full-time employed fathers[12], we see in Figure 1 that less than three percent of the fathers whose child was born prior to 1993 utilized parental leave. After the paternity quota was introduced in 1993, about 30 percent of fathers made use of their right to paternity leave, increasing to 51 percent in 1994 and 59 percent in 1995. More than 70 percent of full-time employed fathers of children born in 2000 took paternity leave.[13] As Figure 1 reveals, the paternity quota had low uptake the first years after implementation, particularly for children born in 1993 and 1994. We will consequently refer to the fathers of these two cohorts as treated in the phase-in-period.

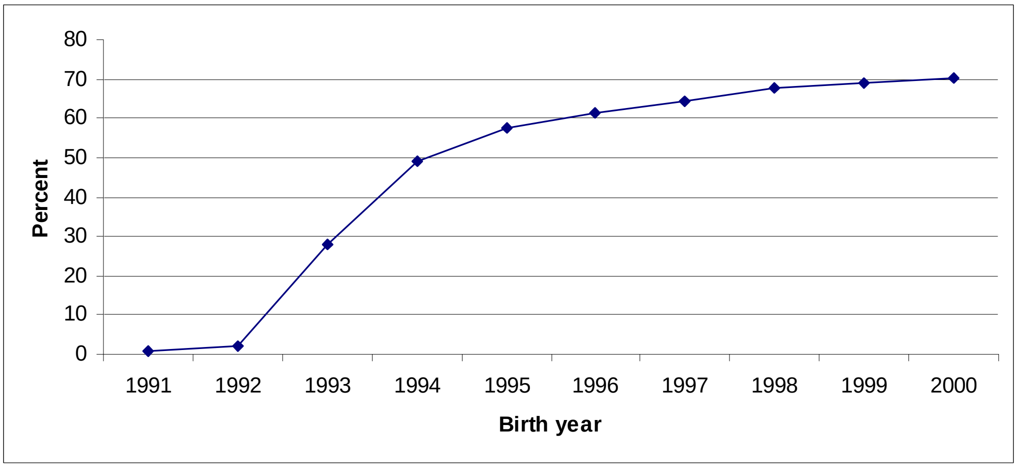

Figure 1: Uptake rates: Percent of fathers in our analytical sample taking paternity leave by birth year of child.

Fathers were entitled to utilize their right to paternity leave up until the child turned three years old. However, 95 percent of all fathers who utilized their right to paternity leave

[10] Fathers have been eligible for parental leave since 1978.

[11] After 2000, a father's income compensation was only reduced if the mother worked less than 75 percent of full-time prior to birth. From 2005, a father's income compensation is independent of how much the mother works prior to birth, but is contingent on the mother being occupationally active while he is on leave.

[12] For a full description of our analytical sample, see Section 3.

[13] Official uptake rates from the Norwegian National Insurance Administration are higher since they calculate the uptake rate of *entitled* fathers.

took leave in conjunction with the mothers' leave during the child's first year. Among fathers taking paternity leave, around 70 percent were on leave for four weeks, 20 percent took less leave, and the remaining 10 percent took more than the designated four weeks of leave.[14] This picture remained relatively constant during our period of study.

We will utilize the introduction of the paternity quota to investigate a causal effect of paternity leave on father involvement. The shadings in Figure 2 illustrate the nature of the experiment. Notably, we construct our experiment based on the age of the youngest child. This is because the father of a child born prior to the introduction of the paternity quota may still be treated if the father is on paternity leave with a younger child. Each row in Figure 2 represents the age of the father's youngest child, and each column represents a given year. To illustrate, the single cell 1997/3 represents fathers whose youngest child turned 3 in 1997. Fathers of each cohort enter into multiple cells diagonally in the figure, according to the age of the father's youngest child. Darkly shaded cells represent fathers treated by the reform after the phase-in-period. These are fathers whose youngest child is born after 1994. At this point nearly 60 percent of the fathers utilized parental leave. Lightly shaded cells represent fathers treated by the reform during the phase-in-period in 1993 or 1994. White cells represent non-treated fathers.

	1992	**1993**	**1994**	**1995**	**1996**	**1997**	**1998**	**1999**	**2000**
1 year									
2 years									
3 years									
4 years									
5 years									
6 years									
7 years									
8 years									

Figure 2: Nature of the experimental design. Darkly shaded cells represent fathers treated by the reform after the phase-in-period, lightly shaded cells represent fathers treated by the reform during the phase-in-period, and white cells represent non-treated fathers.

In addition to the paternity quota, Norway implemented several work-family related policies during our period of study. These policies may have affected mothers' and fathers' long-term involvement. In particular, there was a large extension in paid parental leave between 1986 and 1993. In 1986, Norwegian parents were granted 18 weeks of paid parental leave, but during subsequent years leave rights were gradually extended to 35 weeks in 1992,

[14] Numbers obtained from the Norwegian Labour and Welfare Administration.

and finally to 42 weeks in 1993. Figure 3 shows how many weeks of paid parental leave parents of different age cohorts were granted. Similarly to Figure 2, each cell represents parents of children of a given age in a given year, and parents of a given cohort of children can be followed diagonally in the figure. The figure shows large extensions in parental leave rights prior to the introduction of the paternity quota in 1993. Fathers' *direct* utilization of these extensions was, however, negligible. As discussed above, less than three percent of fathers took paternity leave prior to 1993 and among those taking leave after 1993, only 10 percent took more than the designated four weeks.

	1992	**1993**	**1994**	**1995**	**1996**	**1997**	**1998**	**1999**	**2000**
1 year	32	35	42	42	42	42	42	42	42
2 years	28	32	35	42	42	42	42	42	42
3 years	24	28	32	35	42	42	42	42	42
4 years	22	24	28	32	35	42	42	42	42
5 years	20	22	24	28	32	35	42	42	42
6 years	18	20	22	24	28	32	35	42	42
7 years	18	18	20	22	24	28	32	35	42
8 years	18	18	18	20	22	24	28	32	35

Figure 3: Number of weeks of total parental leave with 100 percent coverage.

In addition to the extensions in paid parental leave, in 1995 both parents became entitled to job protection during one additional year of unpaid leave. In line with paternity leave prior to 1993, few fathers utilized this right.[15] Moreover, in 1998 a cash-for-care subsidy was introduced for families with one or two year old children that did not utilize governmentally subsidized daycare. The cash-for-care subsidy was a tax-free transfer, and, at the time it was introduced, was equivalent to approximately $600 per month. Nearly 80 percent of all families with a one or two year old received the subsidy. The cash-for-care subsidy decreased eligible mothers' labor force participation by 5-6 percentage points, but had no effect on fathers' labor force participation (Schøne 2004; Drange and Rege 2010).

In summary, even if Norway implemented several work-family related policies in addition to the paternity quota, fathers' direct utilization of these reforms has been negligible. Notably, however, if any of these reforms *decreased* mothers' future labor supply, then this may have indirectly motivated fathers to *increase* their labor supply in order to compensate for the family's income loss or because he is less needed in household production (Becker

[15] Only 5 percent of fathers of children born in 2007 utilized their right to unpaid leave. The majority of these (54 percent) were on unpaid leave for two weeks or less (Grambo and Myklebø 2009). Corresponding numbers for 1995-2000 are not available.

1985). In this way, mothers' utilization of extended leave rights may have had a negative impact on father involvement. Consequently, our empirical analyses will investigate whether our estimated effect of the paternity quota is downward biased by these other policy reforms.

3. Data and Sample Description

In our main empirical investigation we will analyze how the introduction of the paternity quota affected fathers' earnings. If the paternity quota increased fathers' long-term involvement, then we should expect a reduction in fathers' long-term earnings, as fathers shift time and effort from market to home production (Becker 1985). We will investigate this hypothesis by utilizing a combination of several official Norwegian registers, prepared and provided by Statistics Norway. The dataset contains records for every Norwegian from 1992 to 2002. The variables include individual demographic information (gender, age, marital status, number of children, birthdates of children), socio-economic data (years of education and earnings, and municipality of residence) and current employment status (full-time, part-time, minor part-time, self-employed).

We restrict our sample to all fathers whose *youngest* child was between 1 and 8 years old during the years 1992 to 2000. Constructing our sample based on the age of the youngest child is important because fathers of children born prior to the introduction of the paternity quota may still be treated if they were on paternity leave with a younger child. The purpose of the remaining sample restrictions is to exclude fathers who are not eligible for paternity leave because of a weak attachment to the labor force First, we limit our sample to fathers who are currently full-time employed.[16] Our definition of full-time employed allows for considerable variation in working hours. According to the most recent data on men's labor force participation, 10 percent of all full-time working men work 30-36 hours per week, 75 percent work 37-43 hours per week and 15 percent work more than 44 hours per week.[17]

Second, since students have a weak attachment to the labor force, we restrict the sample to couples where both parents were older than 25 years when the child was born. Third, we limit our sample to individuals born in Norway to Norwegian-born parents since immigrants in general have substantially weaker labor force attachment (Olsen 2008) and thereby less entitlement to parental leave. Ideally, we would exclude separated couples since

[16] A worker is recorded as full-time employed if he is registered as full-time employed (at least 30 hours work per week) at the end of the year and had earnings above an indexed minimum of about €19 000 in 2010 (2 times "Basic Amount"). We add the earnings restriction because firms are often late in reporting changes in employment status after a work spell has ended.

[17] Labour Force Survey, Statistics Norway, 2010.

fathers not living together with the child's mother are exempt from the paternity quota. However, marital status is potentially endogenous to the reform and we do not observe marital status prior to 1992. Among the fathers in our sample, 91 percent are living together with the child's mother.

Notably, the full-time employment sample restriction may be endogenous if the reform had an impact on the fathers' decisions to be full-time employed. We carefully investigate such possible endogeneities in our data analyses. Clearly, the first best solution would have been to limit our sample to fathers who were full-time employed at the time of the child's birth. However, since we do not observe employment status prior to 1992, we are restricted to using current employment status instead.

The sample selection criteria leave us with a total of 1 127 093 observations for 261 324 fathers of 327 893 children. Notably, in our sample we have several earnings observations for each father. For example, a father with a six year old child in 1992 will have a seven year old child in 1993, and an eight year old in 1994. Consequently, we will observe his earnings in 1992-1994 (See Figure 2, a father is followed diagonally). After 1994 his child is too old to be included in the sample and we do not observe his earnings. However, if this father has a new child in 1995, he will again enter our sample with a one year old in 1996, and a two year old in 1997, etc. Consequently, we will observe earnings for this father in all years except 1995. We use Stata-cluster estimation to correct for multiple observations for each father.

Our data allows us to construct several variables capturing important child, father and mother characteristics. Similarly to employment status, we do not observe pre-birth characteristics for fathers of children born prior to 1993 and consequently we construct our covariates from current characteristics, observed in the same year as we observe outcome. We therefore limit covariates to characteristics that are most likely exogenous to the reform. Moreover, our empirical analyses assure that our results are robust to the inclusion and exclusion of different covariates.

Our empirical analysis utilizes the following covariates, in addition to year fixed effects;

- *Youngest Child Characteristics*: number of older siblings (0,1,2....6, >7) [18], child's age (1,2,…,8), child's gender, birth month (1,2…12)

[18] Parenthetical documentation on any control variable indicates the ranges of the series of categorical variables which characterize the specific trait.

- *Father and Mother Characteristics:* age at birth of youngest child (linear and quadratic), age at birth of first child (linear and quadratic) and education level (not completed high school, high school degree, university degree).[19]

Summary statistics of all observations of fathers in our sample are presented in Table 1. Fathers in our sample were on average 34 years old at the time when the child was born. About 9 percent of the fathers in our sample have not completed a high school degree, and 32 percent have a university degree. The fathers have on average 2.3 children.

In Table 2 we present cohort specific summary statistics for fathers of all children in our sample. In Panel A characteristics are measured one year prior to the child's birth[20] and in Panel B characteristics are measured when the child is three years old. In both panels each father is only observed once for each child. Some cells have missing numbers because data is not available. Notably, with the naked eye we cannot observe any discontinuity in characteristics occurring for fathers of the cohort born in 1995, the first fully treated cohort. Neither can we observe any discontinuity in fathers' earnings measured when the child is three years old.

4. Empirical strategy

We identify the effect on earnings of being on paternity leave by exploiting variation in exposure across fathers over time and the youngest child's age in a difference-in-difference (DD) approach.[21] More specifically, we look at the difference in earnings in a given year between treated and non-treated fathers. However, non-treated and treated fathers in a given year have children of different ages, which alone is likely to have an impact on earnings. To control for such an age effect, we compare the earnings difference to a corresponding earnings difference in a year prior to the introduction of the paternity quota. The deviation between these two differences is attributed to the paternity quota. The identifying assumption is that *absent* the reform, time trends in earnings would be similar for fathers of children of various ages.

In order to utilize the extensive dataset available and to illustrate that our effect estimates are robust to choice of treatment and comparison group, we estimate variation in earnings for all fathers in our sample during the whole period based on the DD-approach

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[19] Education level is potentially endogenous to the reform. However, less than 1 percent of the fathers in our sample reach a higher education level during our period of study.
[20] Since we do not observe pre-birth characteristics for fathers of children born prior to 1993, data prior to birth are not included in our analyses but displayed here for the sake of comparison.
[21] Since the paternity quota had low uptake the two first years after implementation, , it would not be possible to identify any discontinuity in fathers' earnings associated with the introduction date of the paternity quota.

described above: We estimate the incremental effect on earnings of being a father of a child in a certain age in a specific year (i.e. being a father in a specific cell in Figure 2), compared to a common reference group, when time and age trends are controlled for by the inclusion of year and age fixed effects. The reference group is chosen to be seven and eight years old in 1992. Year 1992 was the first year of observations in our data set. Moreover, children seven and eight years old are non-treated during the entire period we observe the individuals.[22]

Our DD-estimates take the following form:

(1) $$\eta_{a,y} = (I_{a,y} - I_{7-8,y}) - (I_{a,1992} - I_{7-8,1992}) \qquad \text{where } y = 1993,1994\ldots.2000 \quad a = 1,2\ldots..6$$

The term ($I_{a,y} - I_{7\text{-}8,y}$) measures in a given year, y, the difference in earnings of fathers of children aged 7-8 and children aged a. The term ($I_{a,1992} - I_{7\text{-}8,1992}$) measures the corresponding difference, measured in 1992. If treated fathers earn less (more) than non-treated fathers, our DD-estimates, η_{ay}, for fathers of children born after the reform will be negative (positive).

In order to estimate the DD-coefficients, η_{ay} , we specify the following regression:

(2) $$I_{iay} = \alpha + \gamma_y Y_y + \delta_a A_a + \eta_{ay}(Y_y \times A_a) + \beta X_{iy} + \varepsilon_{iay}$$

where I_{iay} denotes log-earnings for father i of a (youngest) child aged a (a=1,2…6) in year y (y=1993,1994…2000). Y_y and A_a are vectors with year and age dummy variables, where γ_y and δ_a capture year and age fixed effects. X_{iy} is a vector of father, mother and child characteristics described in Section 3.

The coefficients of interest in Equation 2 are captured by the matrix η_{ay}, which measures the incremental change in earnings for fathers of children of a given age, a, in a given year, y, compared to fathers of seven and eight year olds in 1992. Importantly, if the paternity quota had a negative effect on fathers' earnings, we should be able to identify a pattern associated with treated or non-treated fathers in the estimates of η_{ay} . This pattern should look similar to the step-wise pattern illustrated in Figure 2. We should see significant negative coefficients for each η_{ay} that correspond to treated cells (darkly shaded cells in Figure

[22] An exception is fathers of children 7 years old in 2000. These children were born in 1993 and the fathers are consequently partly treated. This may raise some scepticism for the 2000 estimates. However, we see no effect on this cohort prior to year 2000 (See results in Table 3). Consequently, we consider it unproblematic to use 7 and 8 year olds in the reference group for the year 2000 estimates. Furthermore, a specification test (not reported here) when only 8 year olds constitute the comparison group gives similar but less precise results.

2). Moreover, coefficients for each η_{ay} that corresponds to non-treated cells should not be significantly different from zero (cells with no shading in Figure 2). Significant coefficients in the non-treated will be a violation of our identifying assumption; that time trends in earnings are similar for fathers of children of various ages *absent* the reform.

Even if the estimated coefficients in the non-treated cells are non-significant, our research design may still generate biased estimates if there are unobservable changes in characteristics that are discontinuous, child-cohort specific and occur at the time of implementation of the paternity quota and have an effect on earnings. One possible concern is, for example, that the reform induced couples to have children at a younger age.[23] Then the decrease in earnings among treated fathers may simply be because our treated fathers are of a younger age. We investigate such possible sources of bias in a specification analysis exploring how our estimates are sensitive to included covariates.

Since not all fathers utilized the opportunity to take paternity leave, the treatment is only intentional (Intention To Treat, ITT). In order to capture the effect on fathers who are actually taking paternity leave, we calculate the treatment of the treated (TOT) estimates:

$$(3) \qquad \eta_{ay}^{TOT} = \frac{\eta_{ay}}{\upsilon_{y-a}}$$

where η_{ay}^{TOT} is the treatment of the treated (TOT) effect for fathers of children aged *a* in year *y*, η_{ay} is our estimated treatment effect (ITT) from Equation (2), and υ_{y-a} is the uptake rate for fathers of children born in year *y-a*.[24]

5. Results

5.1. Main Results

Table 3 presents OLS estimates of the DD-coefficients (η_{ay}). Standard errors (in parentheses) are corrected for heteroscedasticity and non-independence of residuals across fathers' earnings observed at different points in time.[25] Year and age fixed effects, as well as relevant control variables for parents and child (Described in Section 3), are all included in the model.

[23] A number of studies find that family policies affect fertility patterns. See Gauthier (2007) for a recent review.

[24] The TOT-estimates are somewhat underestimated since there was a certain uptake of paternity leave also in the comparison group.

[25] Using the "robust cluster(.)" option in Stata.

The table reveals a step-wise pattern in incremental effects on log-earnings for treated fathers consistent with the shading in Figure 2.[26] In particular, we can see that the DD-coefficients of children born after 1994 (treated children) are significant and negative in all years and for all ages of the child. The DD-coefficients for fathers of children born in 1993 or 1994 (treated during Phase-in-Period) are negative, but small and only significant when the child is 1-3 years old, which corresponds well with the phase-in-period of the uptake documented in Figure 2. Importantly, apart for two year olds in 1994, the DD-coefficients are small and not significantly different from zero for children born prior to 1993. This finding is consistent with our identifying assumption that time trends in earnings are similar for fathers of children of various ages *absent* the reform.

We can see that for a father of a given cohort, the treatment effect decreases somewhat as the child gets older, i.e. diagonally in the matrix, but is still significant when the child is five years old.[27] Larger incremental earnings drop for fathers of younger cohorts can largely be explained by the increase in uptake of the reform. Adjusting for this, the earnings drop remains fairly stable across cohorts.

As discussed in Section 2, Norway introduced several extensions in the parental leave legislation during our period of study. Even if fathers' utilization of these extensions were limited, fathers were indirectly affected if mothers' reduced labor supply motivated fathers to increase theirs. If fathers responded to the general extension of the parental leave in 1993 from 35 to 42 weeks (see Figure 3) by increasing their labor supply, then our treatment effects are under-estimated. Note, however, that we find no evidence in Table 3 for fathers responding to the gradual extensions from 18 to 42 weeks parental leave prior to 1993. In contrast, we find that time trends in earnings are similar for fathers of children of various ages. Thus, since fathers' earnings have not been affected by general extensions in parental leave rights prior to 1993, a response to the 1993-extention in general leave rights is unlikely.

In Section 2 we also discussed how the introduction of a cash-for-care subsidy in 1998 had a substantial impact on mothers' but no effect on fathers' labor supply. Consistent with Drange and Rege (2010), Table 3 suggests that the cash-for-care subsidy had no effect on fathers' labor force participation. If the subsidy had an effect, then we would expect to see a change in the DD-coefficients for the fathers of one and two year old children starting in 1998.

[26] When earnings are measured linearly we find the same pattern. We also find the same pattern when 1993 is the reference year, rather than 1992.
[27] Note that the treatment effect for fathers of one and two year olds can partly be explained for some fathers by less than 100 percent earnings compensation when being on leave, see footnote 6.

Altogether, Table 3 provides substantial evidence that the paternity quota had a significant negative long-term effect on fathers' earnings. The effect persists up until our last point of observation when the child is five years old. The incremental effects on earnings for treated fathers lie in the range of 1 to 2.7 percent, suggesting that fathers on average earn 1 to 2.7 percent less as a direct consequence of the paternity quota. When adjusting the ITT estimate for relevant uptake rates, the TOT effect on earnings ranges from 1.6 to 4.5 percent. As a comparison, estimated effects on earnings of an additional year of education are normally in the range of 5 to 10 percent.[28]

5.2. Specification Analysis

The identifying assumption in our DD-approach is that time trends in earnings for fathers of children of various ages would have been similar absent the reform. The fact that we do not observe significant DD-effects on earnings prior to the reform in Table 3 (apart from two year olds in 1994) supports our identifying assumption. However, our estimates may still be biased by changes in characteristics that are discontinuous, child-cohort specific, occurred at the time of implementation of the paternity quota and had an effect on earnings. In Table 2, even if we cannot observe with the naked eye any cohort specific and discontinuous changes in characteristics occurring at the time of the introduction of the paternity quota, we still investigate such possible sources of bias by exploring how our estimates are sensitive to the inclusion of different covariates and different sample restrictions.

We carry out our specification analyses by collapsing all treatment variables of fathers of children born after 1994 (after the phase-in-period) to one treatment variable, and all the treatment variables of fathers of children born in 1993 and 1994 (during the phase-in-period) to one phase-in-treatment variable. The comparison group consists of fathers of children born before the paternity quota was introduced in 1993. Figure 2 illustrates the nature of the experiment: darkly shaded cells are collapsed to form the treatment group, and white cells are collapsed to form the comparison group. Lightly shaded cells represent those treated during the phase-in-period.

The results are reported in Table 4. All models include year and age fixed effects. Models 1-5 add covariates stepwise for child, mother and father characteristics, and

[28] See Cahuc and Zylberberg (2004) for an overview of empirical findings.

municipality fixed effects. We can see that the additional covariates increase the explanatory power of our model (adjusted R-square). However, the treatment estimates remain at around 1.3 percent across the different model specifications. This suggests that the treatment effect is not biased by any cohort specific and discontinuous changes in observable characteristics. The corresponding TOT-estimate ranges from 2.0 to 2.3 percent. Fathers treated in the phase-in-period face a 0.5 percent decrease in earnings on average.

Models 6 and 7 investigate how the treatment estimate is affected by different sample restrictions. In Model 6 we relax the age restriction that both parents should be older than 25 years old when the child was born. When including all parents older than 21 years old, the estimated treatment effect drops to 1.0 percent. Notably, the uptake rate is lower when we relax the age restriction, partly explaining the lower treatment effect. In Model 7 we can see that when tightening the age restriction to parents who were older than 27 years old when the child was born, the estimated treatment effect increases to 1.7 percent.

One possible concern is that the paternity quota affected fertility. In particular, if the reform increases father involvement, this may motivate couples to have another child which they otherwise would not have had. This, in turn, could have an impact on our estimates of long-term treatment effects, since a selected sample of fathers of older children will exit our sample and enter with a younger child. We address this concern in Model 8 by restricting our sample to fathers of single children. The estimated treatment effect remains basically the same.

We have limited our sample to full-time employed fathers. As discussed in Section 2, this restriction is problematic if the reform had an impact on the fathers' decision to be full-time employed. We investigate this assertion in Table 5. In this table we have dropped the sample restriction of full-time employment and the dependent variable is a dummy indicating whether the father is full-time employed. Apart from these changes, Models 1 and 2 correspond to Models 1 and 4 in Table 4. We can see that in both specifications there is a small and insignificant relationship between the treatment variables and full-time employment.[29] This is consistent with the hypothesis that the reform did not have an effect on the fathers' decision to be full-time employed.

5.3. Subsample Analyses: Father's Education Level

[29] Analyzing this relationship within the same research design as Table 3, we find no pattern in the probability of being full-time employed that could be related to the introduction of the paternity quota (table not reported).

In Table 6 we investigate how the response to the paternity quota varies across different levels of education. We utilize the same collapsed-form specification as Model 4 in Table 4. Since uptake rates are likely to vary between sub-groups, we also report the corresponding TOT-estimates.[30] Model 1-Model 3 in Table 6 show substantial differences in response to the paternity quota for different levels of education. The drop in earnings is larger for fathers who have not completed high school. Adjusting for relevant uptake rates amplifies the differences and gives us a TOT-effect of a 3.4 percent drop in earnings for fathers who have not completed high school, compared to 2.4 percent for high school graduates and 1.0 percent for university graduates. The effect for university graduates is not statistically significant. Some studies suggest that lower educated fathers are less involved with their children (Yeung et.al 2001) and our findings may reflect that the paternity quota has a stronger effect on this group where the potential increase in involvement is largest. Alternatively, our findings may reflect that highly educated fathers have a higher opportunity cost of spending more time at home, and are consequently less responsive to the paternity quota.[31]

5.4. Effects on Mothers' Labor Supply

In Table 7 we investigate how the paternity quota affected mothers' labor market participation. Since many mothers do not work or work part-time, marginal changes in mothers' earnings are not a good measure of mothers' labor market responses. Instead we investigate how the reform affected the mothers' likelihood of working. Our analytical sample is the spouses of the fathers in our main analysis. A mother is coded as employed in a given year if she is registered at year end as employed with at least 20 hours per week.[32] Apart from the dependent variable, the analysis is designed in accordance with the analysis reported in Table 3.

The DD-coefficients in Table 7 do not show a stepwise pattern that corresponds to the changes in fathers' earnings reported in Table 3. We can see a strong decrease in labor supply for mothers of one year old children in 1995, most likely due to the extended job protection

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[30] Table A1 in the Appendix reports correlations between background characteristics and the probability of taking paternity leave.

[31] Empirical findings on the association between education level and father involvement is non-conclusive. See e.g. Yeung et.al 2001 for an overview of the literature.

[32] In addition, earnings have to be above an indexed minimum of about €19 000 in 2010 (2 times "Basic Amount"). We add the earnings restriction because firms are often late in reporting changes in employment status after a work spell has ended.

implemented the same year. As expected, the table also shows that the cash-for-care subsidy implemented in 1998 decreased the labor supply of women with one year old children (from 1998) and two year old children (from 1999).[33] However, we cannot see that the paternity quota affected mothers' labor supply.

6. Understanding the Earnings Drop: Time Use Data

The negative effect of paternity leave on long-term earnings is consistent with increased long-term father involvement and a redirection of effort from market to home production. However, there are at least two other stories for why paternity leave could affect fathers' earnings. First, taking time off from work to be on paternity leave may serve as a signal of being more family-oriented rather than career-oriented. Employers may consider such employees as being less devoted and reliable, thus reducing the likelihood of their giving promotions and pay raises. Second, the negative effect on earnings may reflect foregone human capital accumulation while being on leave. The signaling story does not seem plausible, however, because the uptake of the reform was very high within a few years. Moreover, four weeks[34] of foregone human capital accumulation seems unlikely to have an impact on earnings four years later. Nevertheless, more direct evidence for the effect of the paternity quota on father involvement would strengthen the hypothesis of a causal relationship.

Lack of data on hourly wage rates and number of hours worked limits our possibilities to investigate alternative mechanisms utilizing register data. Instead, we turn to data from the Norwegian Time Use Surveys in order to provide more direct evidence for the effect of paternity leave on father involvement.

6.1 Data

The analysis is based on respondent-reported time diaries data from the 1990 and 2000 Norwegian Time Use Surveys. In each of these surveys a representative cross-section sample of the Norwegian population was asked to keep a time diary for two consecutive days (48 hours). In 2000, the diaries were split into 10 minute slots, and in 1990 into 30 minute slots. For each time slot, respondents were asked to report their main activities, where they were at the time, and together with whom. Each respondent was also interviewed to collect

[33] The first fully treated mothers of the cash-for-care subsidy are those giving birth after July 1998. This explains the gradually increasing treatment effect. See e.g. Schone (2004).
[34] As noted in Section 2, 90 percent of leave-taking fathers were on leave for four weeks or less.

demographic and socio-economic background information such as household composition and work hours. Finally, information on respondents' education level and earnings was collected from official registers.

The net sample from the 2000 survey comprised around 3500 individuals, after a response rate of about 50 percent. The corresponding numbers for the 1990 survey was 3000 individuals with a response rate of 64 percent. We exclude time diaries kept during weekends. Thereafter, we restrict our analytical sample in accordance with the selection criteria described in Section 3. First, we exclude fathers who were 25 years or younger at the time when the youngest child was born. Second, we only include fathers reporting that they are full-time workers.[35] Finally, we limit our sample to fathers whose *youngest* child was between one and twelve years old during the years 1990 and 2000. Notably, this last restriction is different from the selection criteria in the register analysis which focused on fathers whose youngest child was between one and eight years old. We include children up until age twelve in this analysis in order to get a sufficiently large comparison group.[36] This leaves us with a total sample of 407 fathers, 186 from the 1990 survey and 221 from the 2000 survey.

Fathers of 1-5 year olds in 2000 are coded as treated, and fathers of 6-7 year olds in 2000 are coded as treated in the phase-in-period (See Figure 1). The comparison group consists of fathers of 8-12 year olds in 2000 and all fathers in 1990. Notably, since we only have data from 1990 and 2000, we cannot connect differential time allocation directly to the reform as a treatment effect. Still, for the sake of simplicity, we will refer to fathers of 1-5 year olds in 2000 as "treated".

We explore changes in fathers' time allocation by observing changes in time spent working versus time spent at home. Time spent working includes all time spent on activities related to paid work (working, lunch break and other breaks, and travel time between work and home). Time spent at home is all time reported being at home, nights included. Furthermore, we attempt to explore changes in father involvement by observing changes in time fathers spent with their children. Following Lamb et.al (1987) we measure father involvement along three different dimensions: Availability, responsibility and interaction. We construct three different dependent variables that may capture these dimensions: We measure

[35] Fathers are recorded as "full-time workers" if they report working regularly at least 37.5 hours per week, which is the statutory "full-time work" in Norway. We do not have data on whether the father is self-employed. As in our analysis on register data, we find that full-time working is not significantly affected by the reform (see Table 8).

[36] A sample of fathers of children up to 8 years old left us with a comparison group for 2000 with only 18 observations. When doing the analyses with this small sample, we get similar but less precisely estimated effects.

availability as all time spent together with the youngest child, irrespective of location (may be away from home). Responsibility is measured as time together with the youngest child when the mother is not present, otherwise along the same lines as availability. Interaction is measured as time where childcare is reported as being the primary activity, such as caring for, playing with, talking with or reading to children living in the household.[37]

In summary, this leaves us with five dependent variables of fathers' time allocation and involvement: Time spent working, time spent at home, time spent together with the child, time spent together with the child when the mother is not present and time spent interacting with children. All five variables are measured as minutes per day, calculated as the average over the two diary days. If the paternity quota had an impact on father involvement and made them redirect their time use towards home production rather than market production, then we would expect to find a negative treatment effect on time spent working, and positive treatment effects on time spent at home and together with the child.

We construct control variables capturing child, father and mother characteristics in accordance with those described in Section 3 when data are available:

- *Youngest child characteristics:* Number of older siblings (0,1,2,>3), child's age (1,2,3,4…12), child's gender.
- *Parent characteristics:* Father's age at birth of youngest child (linear and quadratic), father's and mother's education level (not completed high school, highs school degree, university degree).

We also include dummies for which weekdays the father kept the diary in addition to year fixed effects.

6.2 Results

Summary statistics of time use are reported in Table 9. We find that among fathers of one-five year olds, treated fathers (fathers in 2000) spent less time at work than non-treated fathers (fathers in 1990). Less time at work is mirrored by slightly more time spent at home. In particular, we see diverging trends in time use at work and at home for fathers of 1-5 year olds versus fathers of 8-12 year olds. Similarly, while time spent together with the youngest child drops substantially from 1990 to 2000 for fathers of 8-12 year olds, the time use of

[37] Data allow us to identify time interacting with children, where children are together with the father at the given time slot. Since the father may interact with one child while other children also are present, we cannot measure time a father interacts with a specific child. While Availability and Responsibility are measured as minutes per day together with the youngest child, Interaction is measured as minutes spent interacting with all children in the household.

fathers of 1-5 year olds remains fairly constant over the same period. These patterns are consistent with the hypothesis that fathers affected by the paternity quota have redirected more time into home production and are more involved with their children.

A possible concern when inspecting Table 9 is that the diverging trends for fathers of 8-12 year olds and 1-5 year olds are mainly driven by changes in the trends for fathers of 8-12 year olds. The diverging trends may be due to something that happened to fathers of the older children and not due to the paternity quota. We cannot rule out this possibility. However, summary statistics for similar mothers in the Time Use Surveys reported in Table 10 reveal that trends in time use for both mothers of 8-12 year olds *and* mothers of 1-5 year olds follow the same pattern as time changes for fathers of 8-12 year olds. This suggests that diverging trends for fathers of young and old children are not driven by the changes of the fathers of 8-12 year olds. A possible explanation is that there is a general trend where fathers and mothers spend less time with their children, and that the paternity quota reversed this trend for the father.

Similar to the analysis of register data (see Equation 2), we estimate the following DD-model for each of our five dependent variables:

$$(4) \qquad T^u_{iay} = \alpha^u + \beta^u X_{iy} + \gamma^u Y + \delta^u_a A_a + \eta^u_{1-5}(Y \times A_{1-5}) + \eta^u_{6-7}(Y \times A_{6-7}) + \varepsilon^u_{iay}$$

where T^u_{iay} denotes time (minutes) spent on u (work; at home; with youngest child; alone with youngest child; interacting with children) for father i of a child aged a (a=1,2,3…12) in year y (y=1990,2000), Y and A_{1-5} and A_{6-7} capture year (y = 2000) and age dummies, and X_{iy} is a vector of father, mother and child characteristics described in Section 6.1. The coefficient η^u_{1-5} measures the "treatment effects" of the paternity quota.

Table 11, Panel A summarizes the results from the DD-analysis. We find that treated fathers reduce their time spent working by 79 minutes (Model 1b), and increase their time spent at home by 70 minutes per day (Model 2b), but the effects are not statistically significant. Regarding time use at home, we find that treated fathers increase the time spent together with the child by 64 minutes (Model 3b), and by 38 minutes when the mother is not present (Model 4b). Finally, we find that treated fathers interact 8 minutes more per day with the child, but this effect is not statistically significant (Model 5b). When estimating the models without any covariates, the treatment estimates move by only a few minutes (Model 1a-5a), suggesting that the treatment effects are not affected by compositional changes in the treatment and comparison group. In Table 11, Panel B we report results from a DDD-analysis including similar mothers in the comparison group. These results should be interpreted with caution since mothers time use may also be affected by the paternity quota.[38] Compared to the DD-framework, this approach yields stronger treatment effects on fathers' time use with children.

In summary, the evidence in Tables 9 and 11 is consistent with the hypothesis that fathers affected by the paternity quota have redirected time and effort from market production to home production, and they invest more time in their children. Reduced work hours, mirrored by increased time use together with children, seems to be a plausible explanation for why the paternity quota had a negative effect on fathers' earnings in our analysis of register data.

7. Conclusion

In this paper we have investigated the effects of paternity leave on long-term father involvement. We have utilized variation in exposure to the non-transferable paternity quota of the parental leave as exogenous variation in leave taking. We find strong evidence for a drop in earnings associated with the paternity quota: The four weeks of paternity leave during the child's first year decrease fathers' future earnings by 2.1 percent. The drop in earnings is consistent with increased father involvement, as fathers shift time and effort from market to home production. In order to further investigate this hypothesis, we turn to time diaries from the Norwegian Time Use Surveys and demonstrate that affected fathers spent significantly more time together with their children after the paternity quota was implemented. We also find that treated fathers spent less time working, but the effect is not statistically significant. Together with our analysis of registry data, which provides convincing evidence of a causal effect of the paternity quota on earnings, the time use analyses suggest that the paternity quota had an impact on long-term father involvement.

This study is important because it suggests that paternity leave policies potentially have implications for child well-being (Han, Ruhm and Waldfogel 2009). Increasing empirical evidence suggests that the involvement of a father in his children's lives is important for the children's cognitive and socio-emotional outcomes.[39] The policy relevance of this paper is highlighted by the fact that the European Parliament recently adopted a

[38] As summary statistics on mothers suggested, a separate analysis for mothers (not reported here) results in no significant treatment effects on time use with children, and all estimates are negative.
[39] See for example Lamb (2010) and Tamis-Lemonda and Cabrera (2002).

directive stipulating the minimum requirements for parental leave, including a non-transferrable paternity quota of four weeks.[40] Moreover, the Norwegian paternity quota was extended from 6 to 10 weeks in 2009 and further extensions are expected. Our results suggest that paternity leave has the expected positive effect on long-term father involvement. The next step for future studies is to investigate how paternity leave affects child outcomes.

[40] European Union: Council Directive 2010/18/EU. The 27 member countries are to implement the directive in national legislation within two years.

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Tables:

Table 1: Summary statistics.
Means (and standard deviations).

Outcome variables	
Father's earnings (NOK)	311 740 (180 714)
Mother employed	0.508
Control variables	
Father age at birth	33.78 (4.75)
Mother age at birth	31.27 (3.78)
Father age at birth of first child	28.39 (4.76)
Mother age at birth of first child	26.11 (4.35)
Number of children	2.26 (0.96)
Age of child	3.94 (2.29)
Father not completed high school	0.094
Father high school degree	0.585
Father university degree	0.319
Mother not completed high school	0.081
Mother high school degree	0.599
Mother university degree	0.319
N (observations)	1 127 093
N (children)	327 893
N (fathers)	261 324

Table 2: Summary statistics of parents by child's birth year. Means.

Birth year:	1989	1990	1991	1992	1993	1994	1995	1996	1997	1998	1999
Panel A: All variables observed one year prior to birth											
Outcome variables:											
Father's earnings (NOK)	na.	na.	na.	na.	235 277	239 570	244 906	254 747	267 803	283 807	311 577
Mother employed	na.	na.	na.	na.	0.600	0.615	0.621	0.623	0.625	0.647	0.653
Characteristics:											
Father's age	na.	na.	na.	na.	32.44	32.38	32.47	32.47	32.52	32.46	32.56
Mother's age	na.	na.	na.	na.	29.89	29.88	29.95	30.00	30.10	30.11	30.21
Father's age at birth of first child	na.	na.	na.	na.	28.71	28.74	28.84	28.91	29.05	29.16	29.28
Mother's age at birth of first child	na.	na.	na.	na.	26.48	26.57	26.69	26.73	26.86	27.03	27.13
Father not completed high school	na.	na.	na.	na.	0.100	0.091	0.087	0.080	0.072	0.065	0.057
Father high school degree	na.	na.	na.	na.	0.586	0.591	0.593	0.595	0.594	0.589	0.589
Father university degree	na.	na.	na.	na.	0.310	0.316	0.317	0.322	0.332	0.345	0.352
Mother not completed high school	na.	na.	na.	na.	0.073	0.067	0.066	0.059	0.054	0.049	0.044
Mother high school degree	na.	na.	na.	na.	0.603	0.592	0.577	0.569	0.561	0.544	0.530
Mother university degree	na.	na.	na.	na.	0.322	0.339	0.355	0.370	0.383	0.405	0.424
N	na.	na.	na.	na.	25 130	26 189	26 305	26 842	26 333	24 751	22 153
Panel B: All variables observed when child is three years old											
Outcome variables:											
Father's earnings	266 654	272 947	278 479	287 107	300 051	313 909	334 033	350 240	371 837	na.	na.
Mother employed	0.469	0.478	0.506	0.508	0.525	0.540	0.553	0.535	0.508	na.	na.
Characteristics:											
Number of children	2.24	2.24	2.26	2.27	2.26	2.26	2.24	2.24	2.22	na.	na.
Father not completed high school	0.108	0.108	0.105	0.098	0.093	0.088	0.081	0.074	0.067	na.	na.
Father high school degree	0.564	0.566	0.578	0.591	0.594	0.597	0.607	0.615	0.614	na.	na.
Father university degree	0.325	0.323	0.314	0.308	0.310	0.313	0.310	0.309	0.318	na.	na.
Mother not completed high school	0.097	0.089	0.089	0.079	0.070	0.069	0.069	0.059	0.056	na.	na.
Mother high school degree	0.605	0.611	0.605	0.618	0.619	0.609	0.597	0.594	0.587	na.	na.
Mother university degree	0.296	0.298	0.304	0.302	0.309	0.320	0.332	0.346	0.356	na.	na.
N	15 632	16 802	17 030	16 983	17 406	18 062	18 305	17 747	16 695	na.	na.

Table 3: Main results: Incremental effects on fathers' earnings by age of the child and year.

Dependent variable: Fathers' log earnings

Year

Child's age	1992	1993	1994	1995	1996	1997	1998	1999	2000
1	--	-0.003 (0.004)	-0.012*** (0.004)	-0.013*** (0.004)	-0.018*** (0.004)	-0.021*** (0.004)	-0.019*** (0.004)	-0.025*** (0.004)	-0.023*** (0.004)
2	--	-0.006 (0.004)	-0.010** (0.004)	-0.009** (0.004)	-0.014*** (0.004)	-0.024*** (0.005)	-0.016*** (0.005)	-0.021*** (0.004)	-0.030*** (0.005)
3	--	0.003 (0.004)	-0.004 (0.005)	-0.001 (0.005)	-0.002 (0.005)	-0.008* (0.005)	-0.012*** (0.005)	-0.011** (0.005)	-0.016*** (0.005)
4	--	-0.001 (0.004)	-0.003 (0.005)	-0.005 (0.005)	-0.001 (0.005)	-0.003 (0.005)	-0.004 (0.005)	-0.010** (0.005)	-0.011** (0.005)
5	--	0.005 (0.005)	-0.004 (0.006)	0.004 (0.006)	-0.003 (0.005)	-0.002 (0.005)	0.001 (0.005)	-0.004 (0.005)	-0.011** (0.005)
6	--	0.002 (0.006)	-0.001 (0.006)	-0.004 (0.005)	0.003 (0.005)	-0.003 (0.005)	0.002 (0.005)	-0.001 (0.005)	-0.002 (0.005)
7-8	--	--	--	--	--	--	--	--	--

Notes:
No. of observations: 1 127 093. Adjusted $R^2 = 0.210$. Estimates reflect results from single OLS models, adjusted for year fixed effects, child characteristics (birth order, gender, age and birth month) and parent characteristics (education level, age when the child was born and age when first child was born). *, ** and *** denote significance at 10 percent, 5 percent and 1 percent level. Robust standard errors in parentheses, corrected for non-independence of residuals across fathers at different points in time.

Table 4: Specification Tests. Collapsed treatment and comparison group.

Dependent variable: Fathers' log earnings

	Model 1	Model 2	Model 3	Model 4	Model 5	Model 6	Model 7	Model 8
Treated	-0.0128***	-0.0125***	-0.0135***	-0.0142***	-0.0121***	-0.0098***	-0.0166***	-0.0112**
	(0.0029)	(0.0029)	(0.0028)	(0.0027)	(0.0025)	(0.0023)	(0.0031)	(0.0057)
Treated in phase-in-period	-0.0048**	-0.0052**	-0.0047**	-0.0049**	-0.0046**	-0.0038**	-0.0060**	-0.0020
	(0.0024)	(0.0024)	(0.0023)	(0.0022)	(0.0021)	(0.0019)	(0.0026)	(0.0046)
Covariates included:								
Child characteristics		X	X	X	X	X	X	X
Mother characteristics			X	X	X	X	X	X
Father characteristics				X	X	X	X	X
Municipality fixed effects					X			
Uptake rate	0.622	0.622	0.622	0.622	0.622	0.595	0.629	0.685
TOT-effect (treatment of the treated)	0.0206	0.0201	0.0217	0.0228	0.0195	0.0165	0.0264	0.0164
Other sample restriction						Parents > 21 yrs	Parents > 27 yrs	Fathers of first borns
Adjusted R^2	0.093	0.096	0.143	0.210	0.281	0.236	0.200	0.205
N	1 227 093	1 227 093	1 227 093	1 227 093	1 227 093	1 404 670	903 749	232 007

Notes:
Estimates reflect results from single OLS models, adjusted for year fixed effects, child characteristics (birth order, gender, age and birth month) and parent characteristics (education level, age when the child was born and age when first child was born). *, ** and *** denote significance at 10 percent, 5 percent and 1 percent level. Robust standard errors in parentheses, corrected for non-independence of residuals across fathers at different points in time.

Table 5: Effect of reform on father being full-time employed

Dependent variable:
Father working full-time

	Model 1	Model 2
Treated	0.0021	0.0032
	(0.003)	(0.003)
Treated in phase-in-period	0.0006	0.0009
	(0.002)	(0.003)
Covariates included	Yes	No
Adjusted R^2	0.0341	0.0039
Mean	0.740	0.740
N	1 523 798	1 523 798

Notes:
Estimates reflect results from single OLS models, adjusted for year fixed effects, child characteristics (birth order, gender, age and birth month) and parent characteristics (education level, age when the child was born and age when first child was born). Robust standard errors in parentheses, corrected for non-independence of residuals across fathers at different points in time.

Table 6: Subsample analyses: Fathers' education

Dependent variable: Fathers' log earnings			
	Model 1	Model 2	Model 3
Subsample:	Father not completed high school	Father High school degree	Father university degree
Treated (ITT)	-0.0181**	-0.0162***	-0.0069
	(0.007)	(0.003)	(0.005)
Uptake rate	0.525	0.610	0.667
Treated (TOT)	-0.0346	-0.0266	-0.0104
Treated in phase-in-period	-0.0102	-0.007**	0.002
	(0.006)	(0.003)	(0.004)
Adjusted R^2	0.138	0.129	0.113
N	106 182	659 146	359 110

Notes:
Estimates reflect results from single OLS models, adjusted for year fixed effects, child characteristics (birth order, gender, age and birth month) and parent characteristics (education level, age when the child was born and age when first child was born). *, ** and *** denote significance at 10 percent, 5 percent and 1 percent levels. Robust standard errors in parentheses, corrected for non-independence of residuals across fathers at different points in time.

Table 7: Incremental effects on mothers' labor supply by age of the child and year.

Dependent variable: Mother working at least part-time.

Child's age	Year: 1992	1993	1994	1995	1996	1997	1998	1999	2000
1	--	-0.001 (0.006)	0.005 (0.006)	-0.021*** (0.006)	-0.016*** (0.006)	-0.040*** (0.006)	-0.093*** (0.006)	-0.107*** (0.007)	-0.125*** (0.006)
2	--	0.010 (0.006)	0.005 (0.007)	0.005 (0.007)	-0.013** (0.007)	-0.026*** (0.007)	-0.044*** (0.007)	-0.086*** (0.007)	-0.092*** (0.007)
3	--	-0.003 (0.006)	0.006 (0.007)	0.003 (0.007)	-0.002 (0.007)	-0.011 (0.007)	-0.021*** (0.007)	-0.038*** (0.007)	-0.055*** (0.007)
4	--	0.018*** (0.007)	0.008 (0.007)	0.020** (0.008)	0.017** (0.008)	0.019*** (0.007)	0.005 (0.007)	0.000 (0.007)	-0.008 (0.007)
5	--	0.001 (0.007)	0.013 (0.008)	0.007 (0.008)	0.012* (0.007)	0.012* (0.007)	0.011 (0.007)	0.012 (0.007)	0.000 (0.008)
6	--	0.012 (0.008)	0.003 (0.009)	0.015** (0.008)	0.003 (0.008)	0.003 (0.008)	0.005 (0.008)	0.012 (0.008)	0.008 (0.008)
7-8	--	--	--	--	--	--	--	--	--

Notes:
No. of observations: 1 127 093. Adjusted R^2 = 0.089. Mean of dependent variable: 0.508. Estimates reflect results from single OLS models, adjusted for year fixed effects, child characteristics (birth order, gender, age and birth month) and parent characteristics (education level, age when the child was born and age when first child was born). *, ** and *** denote significance at 10 percent, 5 percent and 1 percent levels. Robust standard errors in parentheses, corrected for non-independence of residuals across fathers at different points in time.

Table 8: Effect of reform on father being full-time employed

Dependent variable:
Father working full-time

	Model 1	Model 2
Treated	-0.0522	-0.034
	(0.0823)	(0.0826)
Treated in phase-in-period	0.076	0.0769
	(0.1194)	(0.120)
Covariates included	Yes	No
Adjusted R^2	0.0429	0.0038
Mean	0.80	0.80
N	508	508

Notes:
Estimates reflect results from single OLS models, adjusted for year fixed effects, child characteristics (birth order, gender, age and birth month) and parent characteristics (education level, age when the child was born and age when first child was born). *, ** and *** denote significance at 10 percent, 5 percent and 1 percent levels. Robust standard errors in parentheses, corrected for non-independence of residuals across fathers at different points in time. Source: 1990 and 2000 survey data (Time use) from Statistics Norway

Table 9: Summary statistics of fathers' time use, by age of the child and year. Minutes per day.

	Allocation of time between work and home				------------------- Time use together with children -------------------						
	Time spent at work		Time spent at home		Availability		Responsibility		Interaction		
	1990	2000	1990	2000	1990	2000	1990	2000	1990	2000	N
1-5 years	525.4	475.6	752.8	763.2	229.6	223.3	83.4	88.4	66.4	58.1	236
	(191.1)	(221.6)	(230.6)	(249.9)	(162.5)	(165.0)	(85.9)	(99.6)	(59.0)	(52.9)	
6-7 years	536.9	449.4	755.8	715.6	245.5	209.4	125.5	90.3	44.7	31.1	53
	(204.2)	(252.8)	(142.9)	(314.9)	(190.6)	(152.4)	(91.5)	(103.1)	(46.0)	(31.1)	
8-12 years	483.6	498.4	791.8	763.0	236.3	170.0	85.2	60.3	32.9	19.4	118
	(152.5)	(200.3)	(222.6)	(186.0)	(152.2)	(141.9)	(97.0)	(99.3)	(48.5)	(31.0)	
N	186	221	186	221	186	221	186	221	186	221	407

Notes:
Measures of time use together with children: Availability = minutes per day together with the child. Responsibility = minutes per day together with the child when the mother is not present. Interaction = minutes per day when caring for the child is reported as the main activity. Source: 1990 and 2000 survey data (Time use) from Statistics Norway.

Table 10: Summary statistics of mothers' time use, by age of the child and year. Minutes per day.

	Allocation of time between work and home				------------------- Time use together with children -------------------						
	Time spent at work		Time spent at home		Availability		Responsibility		Interaction		
	1990	2000	1990	2000	1990	2000	1990	2000	1990	2000	N
1-5 years	314.0	360.6	850.8	877.7	418.5	310.8	298.0	203.5	139.5	120.5	114
	(236.4)	(230.7)	(358.1)	(288.6)	(215.3)	(190.3)	(225.5)	(159.9)	(76.2)	(97.6)	
6-7 years	352.5	381.4	993.8	821.2	330.0	291.2	285.0	203.8	136.9	65.7	25
	(239.3)	(192.1)	(263.1)	(239.0)	(169.3)	(195.2)	(210.4)	(154.1)	(57.6)	(35.2)	
8-12 years	317.9	416.3	926.1	854.7	241.4	219.3	133.6	111.0	45.0	34.6	57
	(211.6)	(213.8)	(224.3)	(230.9)	(152.2)	(156.8)	(128.4)	(106.2)	(54.3)	(34.2)	
N	55	141	55	141	55	141	55	141	55	141	196

Notes:
Measures of time use together with children: Availability = minutes per day together with the child. Responsibility = minutes per day together with the child when the mother is not present. Interaction = minutes per day when caring for the child is reported as the main activity. Source: 1990 and 2000 survey data (Time use) from Statistics Norway.

Table 11: Results from time use data.

	Allocation of time between work and home				------------------- Time use together with children -------------------					
Dependent variable:	Model 1: Time spent at work		Model 2: Time spent at home		Model 3: Availability		Model 4: Responsibility		Model 5: Interaction	
	(a)	(b)	(a)	(b)	(a)	(b)	(a)	(b)	(a)	(b)
Panel A:										
Treated (DD)	-77.06	-79.26	65.17	70.36	63.14*	64.01*	30.65	37.58*	9.56	8.33
	(47.4)	(48.8)	(55.0)	(55.5)	(30.3)	(37.1)	(22.4)	(22.0)	(11.35)	(11.5)
R2 adjusted	0.0292	0.0105	0.0130	-0.0019	0.0068	0.0707	0.0078	0.0839	0.1508	0.162
Mean	496.4	496.4	759.2	759.2	217.0	217.0	83.3	83.3	48.0	48.0
N	407	407	407	407	407	407	407	407	407	407
Panel B:										
Treated (DDD)	-39.80	-43.05	-32.77	-15.6	161.77**	185.2***	96.03*	128.7**	17.73	15.46
	(90.7)	(90.8)	(106.3)	(108.3)	(71.7)	(71.4)	(51.5)	(50.6)	(24.6)	(24.8)
R2 adjusted	0.095	0.085	0.026	-0.015	0.085	0.118	0.201	0.249	0.287	0.292
Mean	452.4	452.4	795.5	795.5	244.4	244.4	119.7	119.7	63.0	63.0
N	603	603	603	603	603	603	603	603	603	603
Covariates included:	No	Yes	No	Yes	No	Yes	No	Yes	No	Yes

Notes:

Estimates reflect results from single OLS models with year and age fixed effects. Covariates are number of older siblings, gender, father's age when child was born, parents' education level and weekdays the diaries were kept. *, ** and *** denote significance at 10 percent, 5 percent and 1 percent levels. Standard errors in parentheses. Source: 1990 and 2000 survey data (Time use) from Statistics Norway.

Table A1: Probability for taking paternity leave.

Dependent variable: Father taking paternity leave		
Covariates:	Coefficient	Std. Err.
Birth order:		
2nd birth order	-0.0378***	0.0022
3rd birth order	-0.1312***	0.0032
4th birth order	-0.2140***	0.0047
5th birth order	-0.2843***	0.0078
6th birth order	-0.3455***	0.0141
7th birth order and higher	-0.4893***	0.0182
Parents' age when child was born:		
Father's age	0.0304***	0.0019
Father's age ^2	-0.0004***	0.0000
Mother's age	-0.0083***	0.0031
Mother's age ^2	0.0001**	0.0000
Parents' age when first child was born:		
Father's age	0.0021	0.0015
Father's age ^2	-0.0001***	0.0000
Mother's age	0.0591***	0.0018
Mother's age ^2	-0.0009***	0.0000
Child's gender: Daughter	-0.0022	0.0015
Father's education level:		
High school degree	0.0409***	0.0029
University degree	0.0308***	0.0032
Mother's education level:		
High school degree	0.1260***	0.0032
University degree	0.2269***	0.0035
Birth month:		
February	-0.0055***	0.0038
March	0.0031***	0.0036
April	0.0117***	0.0036
May	0.0211***	0.0036
June	0.0289***	0.0036
July	0.0324***	0.0036
August	0.0398***	0.0037
September	0.0414***	0.0037
October	0.0475***	0.0037
November	0.0397***	0.0038
December	0.0570***	0.0038
Birth year:		
1995	0.0787***	0.0021
1996	0.1111***	0.0022
1997	0.1357***	0.0024
1998	0.1583***	0.0027
1999	0.1708***	0.0035
R^2 adjusted	0.077	
N	404 262	

Notes:
Sample: Fathers of children born after 1993. Estimates reflect results from single OLS model. *, ** and *** denote significance at 10 percent, 5 percent and 1 percent levels.

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