

**Caring Fathers:
Is the Motherhood Wage Penalty Gender Specific?**

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Abstract

In order to motivate fathers to participate more in care taking of small children, the Norwegian government introduced a paternity quota of the parental leave in 1993. In this paper we investigate whether increased care taking of small children has an effect on fathers' earnings as the child ages. We estimate a difference-in-differences model which exploits differences in fathers' exposures to the paternity quota. We find that the paternity quota had a negative effect on fathers' earnings at least until the child turns 5 years old.

1. Introduction

Gender wage differences have received much attention the last decades. An extensive literature suggests that parenthood, followed by strong specialisation in home and market production respectively, explains most of the wage differences: Women face a motherhood wage penalty that may be explained by less human capital accumulation, redirection of effort from market to home production and/or simply discrimination (see Anderson et al 2002, 2003, Waldfogel 1997, 1999, Joshi et al 1999). Moreover, when becoming fathers, men tend to increase their labour supply in order to compensate for the spouse's income loss, and thereby earn a fatherhood wage premium (Lundberg&Rose 2000, 2002, Angrist & Evans 1998, Waldfogel 1998, Glauber 2008).

In Europe, and in particularly Scandinavia, numerous public policy actions have been taken in order to counterbalance such specialization. The motivation has been that society will benefit from women's participation in the labor market, and children will benefit from father's participation in the care taking. The policies appear to be successful. In Norway mothers' labour force participation has increased substantially the last decades and fathers' participation in home production has also increased. Statistics Norway's time use survey suggests that Norwegian fathers of small children spent significantly less time at work and more time at home in 2000 compared to 1990 (Kitterød et.al 2003). However, despite this advancement, Norway still faces a substantial gender wage gap.

So far, most of the public policy has been directed towards women, aiming at facilitating the combination of care giving and labour market participation and promoting women to enter the labour force. Less attention has been given to fathers' role in the family, and motivating fathers to take a stronger part in the upbringing of the children. Stimulating fathers to participate more in care taking for their children was the main intentions when the Norwegian government in 1993 introduced a non-transferable paternity quota of 4 weeks of the total parental leave of 44 weeks. The idea was that the paternity leave would strengthen the ties between father and child, which may have long term effects on father-child attachment.¹

The introduction of the paternity quota provides a unique opportunity to identify a causal effect on father's long term earnings of caring responsibilities. When fathers also play an important role as a parent and not only a breadwinner, is the wage penalty associated with parenthood still gender specific and exclusive to mothers? For the same reasons that studies

¹ Nepomnyaschy and Waldfogel (2007)

have detected a motherhood wage penalty, we should see a fatherhood wage penalty associated with the paternity quota. One reason is that the paternity quota can reduce accumulation of human capital during the father's paternity leave. Alternatively, the paternity quota may strengthen the father's involvement in childrearing and consequently redirect effort from market to home production. And finally, choosing to take time off to be on paternity leave may serve as a signal of being more family oriented than career oriented, and hence a less devoted employee. Introduction of the paternity quota made it far more costly not to conceal such preferences, and may have induced employers to discriminate against such family oriented fathers.

Our analysis utilizes a comprehensive, longitudinal register database containing annual records for every person in Norway. We estimate a difference-in-differences (DD) model which exploits differences in fathers' exposures to the paternity quota. Our variation focuses on differences in exposure among fathers with similar families before and after the introduction of the paternity quota.

We find that the paternity quota had a significant negative effect on father's income when returning to work. The effect persists up until our last point of observation when the child is five years old. We are not able to identify an effect on mother's earnings of the paternity quota. A likely explanation is that mother's response to the paternity quota is more than offset by the strong effects on mother's labour supply of the cash for care subsidy introduced in 1998, and a general extension of parental leave rights introduced simultaneously as the paternity quota. Still, our findings suggest that the paternity quota had a positive effect on gender wage differences.

The structure of the paper is as follows: We begin with an overview of the literature on motherhood wage penalty in section 2. In section 3 we give a brief description of the legislation of the paternity quota and other relevant work-family related reforms, and describe the data in Section 4. In section 5 we outline our empirical strategy for identifying the treatment effect (on fathers' earnings of the paternity quota) and present our results in section 6. In Section 7 we estimate labour supply effects on mothers of the paternity quota. Conclusion and implications of our findings are in Section 8.

2. Motherhood Wage Penalty

An obvious explanation of the motherhood wage penalty is redirection of effort and time use from market to home production. Female labour force participation has increased substantially the last decades, but the number of women working part time is high. Such differences in labour force participation and hours worked are also likely to affect the hourly wage rate. Absence from the labour market due to childbearing and childrearing results in less accumulation of human capital (Waldfogel 1998, Datta Gupta & Smith 2002). Furthermore, work related human capital depreciates when being absent from the labour market, and rents associated with good job matches may be lost. Low skilled workers are less vulnerable to such earnings erosions and the wage penalty is stronger for high-skilled workers (Andersen et al 2002).

However, even after controlling for less job-related experience and human capital, a significant wage penalty remains. A plausible explanation is segregation in the labour market into lower-paying and more family-friendly jobs (Bayard et al (2003)). Furthermore, a number of studies suggest that increased effort in matters not related to work may have a negative effect on productivity at the labour market (Phipps et al 2001, Andersen et al 2002). This is in line with Becker's theory of increased effort into home production as children increase the relative return. Such drop in productivity is, however, difficult to distinguish from discrimination in the labour market, which also may explain some of the wage differential. Since women generally are the main caregiver in families, they may be regarded as less devoted employees and consequently systematically discriminated against men at the labour market. (Budig & England 2001). Albrecht et al (2009) suggests that leave taking may serve as signal of being less committed to work, especially for men. Hence, employers respond by penalizing fathers who choose to take parental leave.

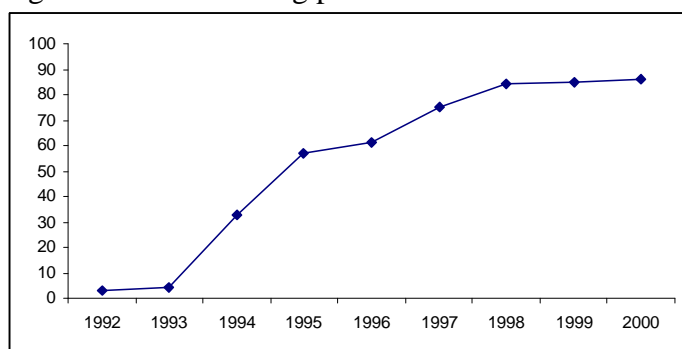
A final explanation to the motherhood wage penalty is related to selection into parenthood: Women may be heterogeneous in an unobserved characteristic such as "career-orientedness" that is both positively correlated with the market wage and negatively correlated with the desire for children. Alternatively, since the market wage rate is an important component in the "cost of children", women with low market productivity may be more likely to become mothers. Several studies find evidence for such negative selection into parenthood (Lundberg and Rose (2000); Korenman and Neumark (1992, 1994))

3. The Paternity Quota

As of April 1st 1993 parental leave rights in Norway were extended by 8 weeks, whereof 4 weeks were reserved exclusively for the father. From then on parents were entitled to 44 weeks leave at full pay or 54 weeks at 80 percent pay after the birth of a child.² 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. Parents can share the parental leave between them as it best suited them, apart from the first six weeks, which are earmarked the mother, and the 4 weeks paternity quota. With few exceptions, the family would lose those 4 weeks of paid parental leave if not taken by the father. Income compensation is based on income to the person being on leave, but father's income compensation was reduced proportionally if the mother did not work full time prior to birth.³

Prior to 1993, less than four percent of all entitled fathers utilized parental leave, see Figure 1. Since the paternity quota was introduced, there has been a sharp increase in the proportion of fathers taking paternity leave. In 1994, 33 percent of all fathers made use of their right to paternity leave, increasing to 57 percent the next year. In 1998 this proportion stabilized at about 85 percent, and in 2004 an estimated 89 percent of all fathers entitled to paternity leave made use of their right.⁴ As Figure 1 reveals, there was a certain phase in period of the uptake of the reform, and it seems unlikely to find any strong effects for fathers of children born prior to 1995.

Figure 1: Fathers taking parental leave. Percent of entitled



Source: The Norwegian Labour and Welfare Administration

² Earnings above 6* "Grunnbeløpet" (around € 50 000 in 2008) are not compensated. Less than 15 percent of all women (aged 17+) earn more than this income ceiling.

³ In 2000 father's income compensation would not be reduced unless the mother worked less than 75 percent of full time prior to birth. From 2005 the calculation of his income compensation is independent of how much the mother worked prior to birth, but is contingent on the mother being occupationally active while he is on leave.

⁴ Fathers are entitled to take paternity leave up until the child turns three years. However, 95 percent of all fathers taking paternity leave take leave in continuation of the mothers' leave.

Among fathers taking paternity leave, around 70 percent are on leave for four weeks. 20 percent take less leave, and the remaining 10 percent takes more than the designated four weeks of leave. This picture remains relatively constant over the period.

We will utilize the introduction of the paternity quota to identify a causal effect on father’s long term earnings of caring responsibilities. The shadings in Figure 2 illustrate the nature of the experiment. Each row represents the age of a child, and each column represents a given year. To illustrate, the single cell 1997/3 represents fathers of children who turn 3 in 1997. Each cohort enters into multiple cells diagonally in the figure. Darkly shaded cells represent fathers that are fully treated by the reform, and white cells represent non-treated fathers. Due to the phase in period of the reform, we define a transition period with partly treated fathers (of children born in 1994 or after after April 1st 1993), represented in Figure 2 with lightly shaded cells.

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

Figure 2: Nature of the experimental design in specification tests

In addition to the paternity quota, other work-family related reforms may have affected mother and father’s attachment to the labour force: First, there was a sharp increase in parental leave rights between 1987 and 1993. In 1986, Norwegian parents were only granted 18 weeks of parental leave, but during subsequent years leave rights were gradually extended to 35 weeks (in 1992). In 1993 parental leave was extended to 42 weeks, including the 4 weeks of paternity quota. Figure 3 shows how many weeks of parental leave parents of different cohorts were granted. Similar to Figure 2, each cell represents fathers (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 diagram.

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

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

The figure shows large deviations in parental leave rights for parents in our sample. However, since few men took paternity leave prior to 1993, these extensions of leave rights are more likely to have affected mother's labour supply.⁵ In addition to extensions in rights for paid parental leave, two additional years of job protection was introduced in 1995. In line with paternity leave prior to 1993, few fathers have utilized this right.⁶ And finally, in 1998 a cash-for-care subsidy was introduced for families with 1 and 2 year old children that were not in day care subsidised by the government. The cash-for-care subsidy was a tax-free transfer, and at the time when it was introduced equivalent to approximately \$600 pr month. Nearly 80 percent of all eligible parent couples receive the subsidy. However, few fathers receive the "cash for care"-subsidy, or take paternity leave (beyond the quota).⁷

While none of these reforms appear to have affected male labour supply, several studies prove significant impact on female labour supply.⁸ Reduced labour supply and earnings for mothers may motivate fathers to increase their labour supply in order to compensate for the family's income loss (Becker 1965). If the paternity quota has a negative effect on father's earnings, his response to other reforms may offset some of the response to the paternity quota. This implies that the estimated reform effect is potentially underestimated. We will perform several specification tests to investigate whether the effect of the paternity is affected by these other reforms.

⁵ Several studies show that increased maternity leave decreases women's future labour market participation. (Ondrich et al. 2003; Lalive and Zweimüller, 2005)

⁶ The Norwegian Labour and Welfare Administration

⁷ The Norwegian Labour and Welfare Administration

⁸ Drange (2009), Schøne (2005)

4. Dataset Description

Our empirical analysis is based on a combination of several official Norwegian registers, prepared and provided by Statistics Norway. The database includes a rich longitudinal dataset containing records for every Norwegian from 1992 to 2002. The variables captured in this dataset include individual demographic information (gender, age, marital status, number of children, birthdates of children), socio-economic data (years of education and income) and current employment status (full time, part time, minor part time, self-employed). Furthermore, data contains personal identifiers for one's spouse, allowing us to link spouses.

We are interested in measuring how the paternity quota affected fathers' future earnings. In our data any person working more than 30 hours a week is recorded as full time employed, and a person working less than 30 hours is recorded as employed part time or minor part time. As we will see in the data analysis, the paternity leave quota did not have an effect on full time employment.⁹ This allows us to focus our analyses on full time employed fathers, because this sample restriction is not endogenous to the reform

Our analytic sample consists of all fathers with the *youngest* child between 1 and 8 years old during the years 1992 to 2000. Focusing our sample selection on the youngest child is important for our identification strategy, since children born before the paternity quota was introduced may still be treated if the fathers were on paternity leave with a younger sibling. We applied the following exclusion criteria to create our final sample of fathers: First, a father who is not living together with his child is not eligible paternity leave. Consequently, we limit our sample to nuclear families at the time of observation: Married¹⁰ couples living together, both biological parents of the child. Second, immigrants may be less responsive to the paternity quota because of weak labor force attachment and different gender-equality traditions as native Norwegian. This leads us to limit our sample to native Norwegians. Third, we include only families we observe during the whole period, which implies that families moving to or from Norway during the observation period are excluded. Fourth, eligibility for paternity quota is conditioned on mother and father's labour force participation prior to child birth. Since students are not eligible for paternity quota, we exclude the youngest parents and restrict the sample to couples where both parents are at least 25 years old when the youngest

⁹ A worker is recorded as full time employed if he is registered as full time employed at the end of the year and had an annual salary above an indexed minimum of about €15 000 in 2008 (2 times "Grunnbeløpet"). We add the salary restriction because firms are often late in reporting changes in employment status after a work spell has ended.

¹⁰ Reliable data on co-habiting couples is not available.

child was born. Finally, as described above, we restrict our sample to fathers employed full time at the time of observation. An important concern is that some of these exclusion restrictions are endogenous to the reform. We perform multiple tests to assure that all exclusion restrictions are exogenous.

This leaves us with a total of 833 873 observations for 199 079 fathers. Notably, in our sample we have several earning observations for each father. For example, a father with a 6 year old child in 1992 will have a 7 year old child in 1993, and an 8 year old in 1994. Consequently, we will observe his income in 1992-1994. After 1994 his child is too old to be included in the sample, and we do not observe his income. However, if this father has a new baby in 1995, he will again enter our sample with a 1 year old in 1996, and a 2 year old in 1997, and so on. Then, we will observe income for this father in all years except from 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 that we include in our regression analyses. We include the following co-variates;

- *Child Characteristics*: number of older siblings (0,1,2,>2)¹¹, child's age (1,2,...,8), child's gender
- *Father and Mother Characteristic*: age at birth of youngest child (linear and quadratic), years of education (<13, 13, 14-15, >16), and indicator for whether municipality of residence is a densely populated or not.

Summary statistics of fathers in our sample are presented in Table 1. The first column presents means and standard deviations of all observations in the sample. In the next three columns of the table we report means and standard deviations for fully treated, partly treated and non-treated fathers respectively. An overall picture is that fully treated, partly treated and non-treated fathers do not differ on observable characteristics. This indicates that fathers do not deviate systematically on unobservable characteristics.¹²

¹¹ Parenthetical documentation on any control variable indicates the ranges of the series of categorical variables which characterize the specific trait.

¹² Systematic deviation based on personal characteristics suggests that fathers are not randomly distributed between the treatment and control group, i.e. there are some (self)selection into the treatment group. Selection based on observable characteristics can be controlled for and is therefore unproblematic. However, selection based on observed characteristics suggests that also unobserved characteristics matter. These can not be controlled for (omitted variables), which may result in biased estimates.

5. Empirical strategy

The true reform effect of the paternity quota is defined as the difference between a treated father's earnings and the earnings he would have made without the reform, i.e. his counterfactual earnings. The fundamental evaluation problem arises because we do not observe his counterfactual earnings. Furthermore, the slow adjustment to the reform suggests that it is unlikely to identify any discontinuity in fathers' earnings associated with the introduction date of the paternity quota (See Figure 1).

We handle these problems by constructing covariate adjusted estimates of the incremental change in earnings for fathers of children of various ages and over different points in time. By exploiting variation in exposure across fathers over time and the child's age, this will provide us with covariate adjusted earnings trends revealing possible patterns related to the paternity quota.

More specifically, based on a difference-in-difference (DD) approach, we first estimate the difference in earnings 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 age effect, we compare this estimated earnings difference to a corresponding earnings difference another year. The deviation between these two difference-estimates is attributed to the paternity quota. Hence, our identifying assumption is that *absent* the reform, time trends in earnings would be similar for fathers of children of various ages. Our rich data set allows us to do several robustness and specification tests addressing this assumption.

In order to utilize the extensive dataset available, we will estimate variation in earnings for all fathers in our sample during the whole period based on the DD-approach 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 the figure), compared to a common reference group, when time and age trends are controlled for. The reference group is chosen to be 7 and 8 years old in 1992, because since 1992 was prior to the introduction of the reform, and 7 and 8 years old are non-treated during the entire period we observe the individuals. Based on the assumption that the difference in earnings between fathers of children in various age groups and fathers of 7-8 years old would have remained constant in the absence of the reform, this will enable us to identify the treatment effects over time and ages of the child.

Our DD estimates take the following form:

$$(1) \quad \eta_{a,y} = (I_{78,y} - I_{a,y}) - (I_{78,1992} - I_{a,1992}) \quad \text{where } y = 1993, 1994, \dots, 2000$$

$$a = 1, 2, \dots, 6$$

The term $(I_{78,y} - I_{a,y})$ measures in a given year, y , the difference in earnings of fathers of children aged 7-8 and children aged a . The term $(I_{78,1992} - I_{a,1992})$ measures the corresponding difference, measured in 1992. This controls for differences in earnings that are related to the child's age. If treated fathers earn less (more) than non-treated fathers, our DD estimates, η_{ay} , for fathers of children born after the reform will turn out negative (positive).

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

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

where I_{iay} denotes log-earnings for father i of a child aged a ($a=1,2,\dots,6$) in year y ($y=1993,1994,\dots,2000$). Y and A are dummy variables, where γ_y and δ_a captures year and age specific effects. X_{iy} is a vector of father, mother and child characteristics described in Section 3.

The matrix of interest is estimates of 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 7-8 years old in 1992. As an illustration, an estimate of $\eta_{1,1997}$ of -0.02 implies that the incremental effect on earnings for fathers of 1 year olds in 1997 compared to fathers of 7-8 years old in 1992, is a 2 percent reduction, when age and time effects are controlled for. Importantly, if the paternity quota had an effect of fathers' earnings, we should be able to identify a pattern associated with treated or non-treated fathers in the estimates of η_{ay} . This patten should look similar to the step-wise pattern illustrated in Figure 2..

6. Results

6.1. Main results

Table 2 presents OLS estimates of the DD-coefficients (η_{ay}). Standard errors (in parenthesis) are corrected for heteroscedasticity and non-independence of residuals across fathers at different points in time. Year and age specific effects, as well as relevant control variables on parents and child, are all included in the model.

The table reveals a clear step-wise pattern in incremental effects on earnings for treated fathers: Treatment effects on earnings of being the father of a child born after 1994 are significant and negative in all years and for all ages of the child, all up to the highest observed age of 5 years (i.e. for fathers of children born in 1995). These strong and statistically significant effects are in sharp contrast to small and insignificant effects for fathers of children born in 1994 and earlier. The incremental effects on earnings for treated fathers lay in the range of 1 to 2.5 percent, suggesting that fathers on average earn 1 to 2.5 percent less as a direct consequence of the paternity quota.

Not-significant treatment effects for fathers of children born in 1993 and 1994 correspond well with data on slow adjustment to the reform, and confirm our assumption that this group is only partly treated. Furthermore, slow adjustment may also explain why incremental effects on earnings increase in strength over time, being stronger for fathers of younger cohorts. Treatment effects decrease as the child gets older but is, as pointed out above, still significant when the child turns 5 years. All together, this gives substantial evidence that after a certain time of adjustment and sufficiently high uptake of the reform, the paternity quota has had a negative impact on fathers' earnings. This suggests that the well documented wage penalty is not gender specific. There is also a wage penalty associated with fathers' care taking of small children.

6.2. Interaction with other reforms

As noted earlier, other work-family related reforms introduced during the same period may have affected fathers' earnings. The sharp increase in parental leave rights from 18 to 42 weeks between 1987 and 1993 (see Figure 3) may have directly affected both mother's and father's labour supply. The figure shows that our control group, 7-8 years old in 1992, had substantially weaker leave rights than the remaining sample. However, low utilization rates among fathers of parental leave prior to the paternity quota suggests that fathers to a little extent responded to these extensions by extending their leave taking. On the other hand, it is not unlikely that reductions in mother's labour supply¹³ motivated fathers to increase theirs. If this is the case, the treatment effect is under-estimated. However, we find no evidence in Table 2 for father's responses to the gradual extended leave rights: The coefficients do not follow a pattern that could be traced back to extensions in leave rights.

¹³ Several studies show that increased maternity leave decreases women's future labour market participation. (Ondrich et al. 2003; Lalive and Zweimüller, 2005)

Another reform with substantial impact on mother's labour supply is the cash for care subsidy introduced in 1998.¹⁴ We could therefore expect that father's earnings increased up until the child is three years old. This may explain why the strength in father's response is somewhat smaller in 1998 than in 1997, and suggests that the treatment effect of the paternity quota is in fact stronger than reported in Table 2.

6.3. Specification tests and robustness checks

Our DD estimates rest on the assumption that earning differences between fathers of children of various ages would have remained constant over time absent the reform. If the treatment group deviates from the control group of other reasons than the reform, and these deviations affect earnings, we may mistakenly interpret this as a treatment effect. Although Table 1 gives us no reason to believe that treated fathers is a selected group, i.e. deviates systematically from non-treated fathers, it is crucial to do a more thorough investigation if this.

There are basically four sources for selection. First, specific characteristics may change due to underlying trends related to fathers of different age groups. Second, the treatment may in itself affect fathers' characteristics. Third, if fertility is affected by the reform, fathers with certain characteristics may be more likely to be exposed to the treatment. And fourth, our sample selection criteria may be endogenous to the reform.

We can only check if selection has taken place based on observable characteristics. Although the econometric models handle and control for selection based on observable characteristics, it is still problematic since selection based on observable characteristics is a strong indicator of selection based on also unobservable characteristics.

We apply a number of specification and robustness tests in order to check whether the treatment estimate is biased due to selection and omitted (unobservable) variables. In order to facilitate these tests, we produce a single estimate of the treatment effect. Based on a similar DD approach, we define the treatment group as fathers that are proven to be treated, i.e. fathers of children born after 1994. The control group consists of fathers of children born before the paternity quota was introduced. Partly treated fathers form a separate group. Figure 2 illustrate the nature of the experiment: Darkly shaded cells are collapsed to form the treatment group, and white cells are collapsed to form the control group. Lightly shaded cells represent partly treated fathers.

¹⁴ Drange (2009) finds that mother's labour supply dropped by XX percent following the reform. Schøne (2005) found a corresponding estimate of XX percent.

Results are in Table 3. The treatment effect is 1.7 percent and implies that treated fathers have on average 1.7 percent lower earnings than non-treated fathers (Table 3, Model 1). Partly treated fathers face a significant wage penalty of 0.8 percent, about half the size. In order to check if the estimate is sensitive to included covariates, we add covariates stepwise. The estimate remains constant at 1.7 percent (Model 2-Model 4), which indicates that treated fathers do not deviate from non-treated fathers.

Finally, we check if the sample selection criteria are exogenous to the reform. As explained above, we exclude fathers not working fulltime in our sample. If the paternity quota induced fathers to work part time rather than full time, we underestimate the treatment effect. Results are reported in Table 4. The reform did not affect whether father worked full time (Model 1). Another sample selection criteria is that parents are older than 25 years when the child was born, in order to exclude student and assure a certain attachment to the labour force. Here we find a significant “treatment effect” which implies that treated fathers are systematically older than non-treated fathers (Model 2 (dummy) and Model 3 (continuous)). This may be due to underlying fertility trends.

In general, the estimated treatment effect will be biased if introduction of the reform affected parents’ fertility pattern through unobserved variables that also determine earnings.¹⁵ This identifying assumption may be difficult to defend. Since we define the treatment group nearly two years after the introduction of the reform, delayed child births are unlikely to affect our estimate. However, several studies find that Scandinavian policies intended to facilitate the combination of work and family obligation has contributed to high fertility rates, *and* to raise employment rates for mothers of small children.¹⁶ If the fertility response to the paternity quota is affected by unobservable characteristics that also affect earnings, our estimated treatment effect will be biased. In fact, studies based in Norwegian register data (Lappegård, 2008) find that parents’ fertility response to the paternity quota depends on education level, age and marital status (married/cohabiting). Still, our summary statistics in Table 1 provide evidence for no such selection mechanisms. Furthermore, fertility trends measured by the total fertility rate during our period of observation indicate that fertility responses are negligible.¹⁷

¹⁵ See footnote 12

¹⁶ For an overview of the literature, see Gauthier (2007).

¹⁷ See Statistics Norway’s homepages for the development in total fertility rate: <http://www.ssb.no/fodte/fig-2008-04-09-01.html>

6.4. Subsample analysis

By doing subsample analyses, we may get a deeper understanding of the mechanisms behind the wage penalty, and which fathers are being “punished” most. It is, however, important to have in mind that, since the estimated treatment effect is an average over all fathers, the treatment estimate is directly affected by the take up rate, which also is expected to vary between subgroups. Furthermore, mechanisms behind the wage penalty cannot be identified solely by doing subsample analysis, since various mechanisms may apply to different subgroups and counterbalance each other. Results from subsample analysis provide therefore far from sufficient evidence to draw certain conclusions.

Results from a number of subsample analyses are presented in Table 6 and Table 7. We find substantial differences in response when taking education level into account. (Table 6) In general: The father’s earnings drop is significantly lower for high school drop-outs than those who completed a high school degree (Model 1), and the opposite applies for different levels of his spouse’s education level (Model 2). The differences are even stronger when creating subgroups based on relative education level between the spouses: Fathers with a lower educated spouse (than him) faces an earnings drop of only 1.2 percent (Model 3). On the other hand, fathers with a higher educated spouse earn a wage penalty of 2.8 percent.¹⁸ It appears implausible that this is related to variations in drop in human capital accumulation or discrimination. Rather, it appears likely that this reflects responsiveness to alter the degree of specialization of responsibilities within the family, and that this effect is especially strong on couples with a more even division of responsibilities for home and market production.¹⁹

For parents of children of first or second birth order, the treatment effect is significantly stronger than for parents of children of higher parities (Table 7, Model 1). The most plausible explanation is again the degree of specialization within the family. Becoming a parent is a milestone. The new set of responsibilities requires that the couple defines the roles as mother and father, and several studies find that the specialization and gender imbalances within the family establishes when couple becomes parents (on Norwegian data, see Kitterød 2002). As children of higher parities are born, the parents are less responsive to policy changes aiming at affecting the division of responsibilities and parent roles. However, since previous studies indicate that fathers of first borns are more likely to actually make use of the

¹⁸ In a subsample analysis not reported here we found that fathers who were high school dropouts and married to a women with more than a high school degree, the wage penalty were nearly 4 percent.

¹⁹ Lundberg and Rose (1999) find that specialization in home versus market activities is in general less pronounced among higher educated parents.

right for paternity leave (Lappegård, 2008), a stronger drop in earnings may simply reflect a higher uptake to the reform among fathers of first and second borns.²⁰

7. Effects on mothers' labour supply

Gender wage equalization is of high political priority, and was among the main intentions when the paternity quota was adapted. Extensions of the paternity quota have recently been suggested²¹ in order to stimulate reduced specialisation in women's and men's home and marked production respectively. We have found evidence for fathers' reduced specialisation in marked production. It is of similar interest to reveal if mothers have responded accordingly, redirecting more effort into the labour market. Furthermore, we have found no evidence for fathers responding to changes in mother's earnings due to public policy reforms that have taken place during the last decades. However, we cannot rule out that other factors affected mother's earnings, and that the treatment effect is due to such other changes. By estimating changes in mother's participation at the labour market in accordance with our identification strategy for the treatment effect of the paternity quota, we may provide sufficient evidence to rule out father's responses other than that of the paternity quota.

Since a significant number of mothers of small children stay at home or work part time, a change in labour supply is likely to be reflected in more mothers entering the labour market, rather than a marginal increase in earnings.

We estimate how the likelihood of mothers participating in the labour market is affected by the paternity quota. Labour market participation is measured as a dichotome variable indicating whether or not the mother is working at least part time. She is registered as at least part time working if she works at least 20 hours a week, registered at year end, and having a salary above a given minimum.²² The remaining design of the analysis is as described in section 4. We find that mothers of children up to the age of 3 years significantly reduce their labour supply compared to the reference group, most likely related to the cash for care subsidy introduced in 1998, and two years job protection introduced in 1995. More importantly, we find no stepwise pattern that corresponds to fathers' earnings (Table 2), which suggests that the treatment effect of the paternity quota is most unlikely induced by changes in mothers' labour supply.

²⁰ Recall that the treatment effect is estimated as an average over all fathers of a certain cohorts in a given year. Higher utilization will, *cet. par.*, lead to a stronger treatment effect.

²¹ St.meld. nr. 1, 2008-2009: The National Budget

²² About €15 000 in 2008 (2 times "Grunnbeløpet").

8. Conclusion

In this paper, we have estimated the effects on fathers' earnings of increased care taking for small children. We have utilized variation of exposure to the non-transferrable paternity-quota of the parental leave as exogenous variation in fathers' care taking of children. We find strong evidence for a fatherhood wage penalty associated with increased care taking for small children: Fathers of children born after the introduction of the paternity quota earn nearly 2 percent less than other fathers at least until the child turns five years old.

Effects on fathers' earnings may reflect changes in hours worked, as well as changes in the wage rate. Furthermore, changes may be supply driven – reduced labour supply and/or productivity among fathers in order to put spend more time on family related activities, or from the demand side – employer discrimination of fathers of small children. Lack of data on hourly wages or hours worked limits our possibilities to identify mechanisms that take place. However, subsample analysis suggests that a plausible explanation for the estimated drop in earnings associated with the paternity quota is related to less specialisation in responsibilities for market and home production respectively.

The implications of our finding are extensive. Our results prove that the well documented “motherhood wage penalty” is indeed not gender specific. Rather, the wage penalty is related to parenting and care taking. Attentive fathers, attentive in terms of participating in their children's upbringing from an early point in time, face a wage penalty in line with mothers.

The intention of the paternity quota was to improve gender balances in the labour market and in home production, as well as to strengthen the ties between fathers and children from an early point in life. Since the paternity quota was introduced in 1993, it has been extended from 4 to 6 weeks (in 2008), and was recently suggested to be extended to 10 weeks. However, up until now no studies have identified any long term effects of the paternity quota. Our findings indicate that gender balances is fact improved, either by reduced labour supply among fathers, or a drop in the wage rate due to a redirection of effort and accumulation of human capital from market production to home production. This is in line with the intentions of the paternity quota, and suggests that extensions of the paternity quota may be adequate actions in order to enhance gender equalisation in the labour market and in the family.

One may claim that although gender wage balances are improved by reduced earnings for fathers, the family as a whole do not benefit from the reform unless his wage drop is counterbalanced by increased female earnings. However, to understand the implications for the family, identification of mechanism is crucial. If father's earnings are reduced due to a

redirection of effort, drawing conclusions of the effects on well being for affected families requires further analyses. First, increased attachment between father and child has an egenreverdi affecting all family members. Furthermore, fathers increased effort in child caring and rearing may have long term effects on children's well being and productivity.

References:

- Albrecht, J.W, P.A. Edin, M. Sundström, S.B. Vronan (1999): "[Career interruptions and subsequent earnings: a reexamination using Swedish data](#)", Journal of Human Resources, 34(2)
- Anderson, D., M. Blinder and K. Krause (2002): "The Motherhood Wage Penalty: Which Mothers Pay It and Why?", The American Economic Review, Vol. 92, No. 2
- Becker, G. (1965): "A theory of the Allocation of Time", The Economic Journal, Vol. 75, No. 299.
- Becker, G. (1985): "Human Capital, Effort, and the Sexual Division of Labour", Journal of Labour Economics, 1985, vol. 3, no. 1.
- Budig, M. and P. England (2001): "The Wage Penalty for Motherhood", American Sociological Review, Vol. 66.
- Drange, N. (2009): "Crowding out of Dad? Labour Supply Responses of the Cash for Care Subsidy", forthcoming.
- Ekberg, J., R. Eriksson and G. Friebel (2005): "Parental Leave – A Policy Evaluation of the Swedish Daddy-Month Reform", IZA DP No. 1617.
- Gauthier, A.H. (2007): "The impact of family policies on fertility in industrialized countries: a review of the literature", Population Research and Policy Review, Vol. 26, No. 3.
- Glauber, R. (2008): "Race and Gender in Families at Work: The Fatherhood Wage Premium", Gender and Society, Vol. 22, No. 1.
- Han, W., C. Ruhm and J. Waldfogel (2007): "Parental Leave Policies and Parents' Employment and Leave-Taking", IZA DP No. 3244.
- Joshi, H., P. Paci and J. Waldfogel (1999): "The Wages of Motherhood: Better or Worse?", Cambridge Journal of Economics, 23.
- Kitterød, R.H. (2002): "Store endringer i småbarnsforeldres dagligliv". Samfunnsspeilet, 4/5.
- Kitterød, R.H. and R. Kjeldstad (2003): "A new father's role? Employment patterns among Norwegian fathers 1991-2001". Economic Survey 1 2003, Statistics Norway.
- Kunze, A. and M. Ejrnaes (2004): "Wage Dips and Drops anround First Birth?", IZA DP No. 1011.
- Kunze, A. and K. Troske (2007): "Comparative Advantage or Discrimination?", IZA DP No. 3052.
- Lappegård, T. (2008): "Changing the Gender Balance in Caring: Fatherhood and the Division of Parental leave in Norway". Population Research and Policy Review, 27.

- Lalive, R. and J. Zweimüller (2005): “Does Parental Leave Affect Fertility and Return-to-Work? Evidence from a True Natural Experiment”, IZA DP No. 1613.
- Lundberg, S. and E. Rose (1999): “The Effects of Sons and Daughters on Men’s Labour Supply and Wages”,
- Lundberg, S. and E. Rose (1999): “The Determinants of Specialisation within Marriage”,
- Lundberg, S. and E. Rose (2000): “Parenthood and the Earnings of married men and women”, *Labour Economics* 7.
- Nepomnyaschy, L. and J. Waldfogel (2007): “Paternity Leave and Fathers’ Involvement with their Young Children”, *Community, Work and Family*, Vol. 10. No. 4.
- Ondrich, J., C.K. Spiess, Q. Yang and G.G. Wagner (2003): “The Liberalization of Maternity Leave Policy and the Return to Work after Childbirth in Germany”, *Review of Economics of the Household*, Vol. 1, No. 1-2
- Phipps, S., Burton, P. & Lethbridge, L. (2001): “In and out of the labour market: longterm income consequences of child-related interruptions to women’s paid work”, *Canadian Journal of Economics*, 34, 411-429.
- Schøne, P. (2005): “The effect of a family policy reform on mother’s pay: A natural experiment approach.” *Review of Economics of the Household*, 3
St.meld. nr. 1, 2008-2009: The National Budget.
- Tanaka, S. and J. Waldfogel (2007): “Effects of Parental Leave and Work Hours on Fathers’ Involvement with their Babies”, *Community, Work and Family*, Vol. 10, No. 4.
- Waldfogel, J. (1997): “The Effect of Children on Women’s Wages”, *American Sociological Review*, Vol. 62.
- Waldfogel, J. (1998): “Understanding the Family Gap in Pay for Women with Children”, *The Journal of Economic Perspectives*, Vol. 12, No. 1

Tables:

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

Variable	All	Fully treated	Partly treated	Non treated
Income father	318 700 (195 300)	350 500 (253 500)	328 100 (216 300)	303 100 (156 200)
Income mother	138 200 (102 000)	154 000 (109 200)	148 400 (102 900)	129 000 (97 700)
Age at birth, father	33.92 (4.61)	33.90 (4.76)	33.96 (4.69)	33.92 (4.53)
Age at birth, mother	31.45 (3.76)	31.43 (3.75)	31.45 (3.77)	31.45 (3.75)
Year of schooling, father	13.52 (2.78)	13.67 (2.86)	13.57 (2.75)	13.45 (2.75)
Years of schooling, mother	13.24 (2.49)	13.73 (2.55)	13.42 (2.44)	12.99 (2.45)
Number of children	2.39 (0.94)	2.30 (0.99)	2.39 (0.97)	2.43 (0.91)
Population density (residence municipality)	0.77 (0.42)	0.79 (0.41)	0.78 (0.42)	0.76 (0.43)
N	833 873	200 832	138 819	494 222
N (fathers)	199 079			

Table 2: Incremental effects on earnings by age of the child and year.

Year:	1992	1993	1994	1995	1996	1997	1998	1999	2000
Age:									
1	--	0,003 (0,004)	-0,008* (0,005)	-0,011** (0,005)	-0,014*** (0,005)	-0,019*** (0,005)	-0,015*** (0,005)	-0,026*** (0,005)	-0,020*** (0,005)
2	--	-0,006 (0,004)	-0,004 (0,005)	-0,006 (0,005)	-0,014*** (0,005)	-0,022*** (0,005)	-0,015*** (0,005)	-0,021*** (0,005)	-0,033*** (0,005)
3	--	0,004 (0,005)	-0,002 (0,005)	0,004 (0,005)	0,000 (0,005)	-0,008 (0,005)	-0,009* (0,005)	-0,011** (0,005)	-0,017*** (0,006)
4	--	-0,003 (0,005)	-0,004 (0,005)	-0,007 (0,006)	0,001 (0,006)	-0,006 (0,005)	-0,008 (0,005)	-0,012** (0,005)	-0,017*** (0,006)
5	--	0,006 (0,005)	-0,003 (0,006)	0,005 (0,006)	-0,003 (0,005)	0,004 (0,005)	0,000 (0,005)	-0,007 (0,006)	-0,015** (0,006)
6	--	0,003 (0,006)	0,002 (0,006)	-0,002 (0,006)	0,003 (0,006)	-0,001 (0,006)	0,006 (0,006)	-0,002 (0,006)	-0,004 (0,006)
7-8	--	--	--	--	--	--	--	--	--

No of observations: 833 873

R-squared = 0.239

Note: *, ** and *** denote significance at 10 percent, 5 percent and 1 percent level. Robust standard errors in parenthesis, corrected for non-independence of residuals across fathers at different points in time.

Table 3: Main results. Collapsed groups

	Model 1	Model 2	Model 3	Model 4
Treated	-0.017*** (0.003)	-0.017*** (0.003)	-0.017*** (0.003)	-0.017*** (0.003)
Partly treated	-0.008*** (0.003)	-0.008*** (0.003)	-0.008*** (0.003)	-0.008*** (0.003)
Co-variates included:				
Child characteristics		X	X	X
Mother characteristics			X	X
Father characteristics				X
R-squared	0.097	0.098	0.170	0.239
N	833 873	833 873	833 873	833 873

Note: *, ** and *** denote significance at 10 percent, 5 percent and 1 percent level. Robust standard errors in parenthesis, corrected for non-independence of residuals across fathers at different points in time.

Table 4: Specification tests (sample selection criteria)

	Model 1:	Model 2:	Model 3:
Dependent variable:	Father working full time	Parents' age at birth > 25	Fathers age (cont.)
Independent variable:			
Treated	-0.003 (0.003)	0.009*** (0.002)	-0.175*** (0.043)
Partly treated	-0.002 (0.003)	0.004 (0.002)	0.148*** (0.035)
Excluded sample criteria:	Father working full time	Parents' age at birth > 25	Parents' age at birth > 25
R-squared	0.065	0.641	
N	1091276	986442	986 442

Note: *, ** and *** denote significance at 10 percent, 5 percent and 1 percent level. Robust standard errors in parenthesis, corrected for non-independence of residuals across fathers at different points in time.

Table 6: Subsample analyses: Father's and mother's education

	Model 1a	Model 1b	Model 2a	Model 2b	Model 3a	Model 3b
Subsample:	Mother edu < father edu	Mother edu > father edu	Father high school drop-out	Father high school degree	Mother high school drop-out	Mother high school degree
Treated	-0.012** (0.005)	-0.028*** (0.005)	-0.024*** (0.005)	-0.015*** (0.004)	-0.014*** (0.004)	-0.020*** (0.004)
Partly treated	-0.006 (0.003)	-0.016*** (0.004)	-0.012*** (0.004)	-0.007** (0.003)	-0.006* (0.004)	-0.010*** (0.003)
R-squared	0.236	0.218	0.149	0.190	0.219	0.213
N	362162	266028	251143	577029	342184	488650

Note: *, ** and *** denote significance at 10 percent, 5 percent and 1 percent level. Robust standard errors in parenthesis, corrected for non-independence of residuals across fathers at different points in time.

Table 7: Subsample analyses

	Model 1a	Model 1b	Model 2a	Model 2b	Model 3a	Model 3b
Subsample:	Birth order < 3	Birth order > = 3	Daughter	Son	Father age at birth 35 -	Father age at birth 35 +
Treated	-0.021*** (0.004)	-0.011** (0.005)	-0.018*** (0.004)	-0.016*** (0.004)		
Partly treated	-0.011*** (0.003)	-0.005 (0.004)	-0.006 (0.004)	-0.011*** (0.004)		
R-squared	0.245	0.232	0.238	0.240		
N	495269	338604	408924	424472		

Note: *, ** and *** denote significance at 10 percent, 5 percent and 1 percent level. Robust standard errors in parenthesis, corrected for non-independence of residuals across fathers at different points in time.

Table 8: Incremental effects on mothers' labour supply by age of the child and year.

Year:	1992	1993	1994	1995	1996	1997	1998	1999	2000
Age:									
1	--	0,000 (0,007)	0,004 (0,007)	-0,015** (0,007)	-0,012 (0,007)	-0,038*** (0,007)	-0,090*** (0,007)	-0,103*** (0,008)	-0,115*** (0,008)
2	--	0,008 (0,007)	-0,001 (0,007)	0,003 (0,007)	-0,009 (0,007)	-0,029*** (0,008)	-0,047*** (0,008)	-0,089*** (0,008)	-0,085*** (0,008)
3	--	-0,011 (0,007)	-0,004 (0,008)	-0,008 (0,008)	-0,008 (0,008)	-0,018** (0,008)	-0,028*** (0,008)	-0,046*** (0,008)	-0,059*** (0,008)
4	--	0,012 (0,007)	-0,001 (0,008)	0,009 (0,009)	0,006 (0,009)	0,009 (0,008)	-0,004 (0,008)	-0,009 (0,008)	-0,018** (0,008)
5	--	-0,001 (0,008)	0,005 (0,009)	-0,003 (0,009)	0,005 (0,008)	0,003 (0,008)	0,004 (0,008)	0,003 (0,008)	-0,009 (0,009)
6	--	0,010 (0,009)	0,000 (0,009)	0,010 (0,009)	-0,002 (0,008)	-0,003 (0,008)	-0,008 (0,008)	0,007 (0,009)	0,001 (0,009)
7-8	--	--	--	--	--	--	--	--	--

No of observations: 833 783

R-squared = 0.089

Note: *, ** and *** denote significance at 10 percent, 5 percent and 1 percent level. Robust standard errors in parenthesis, corrected for non-independence of residuals across fathers at different points in time.