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Trapped at home: The effect of mothers' temporary labor market exits on their subsequent work career ☆

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HIGHLIGHTS

• We investigate how mothers' decision to stay home with a toddler affects their careers
• Identification is based on a policy increasing mothers' incentives to work less
• Earnings and full-time employment is affected even after eligibility expired
• Mothers with low education or low pre-reform earnings drive the effects

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ABSTRACT

This paper investigates how mothers' decision to stay at home with young children affects their subsequent work careers. Identification is based on the introduction of the Cash-for-Care program in Norway in 1998, which increased mothers' incentives to withdraw from the labor market when their child was one and two years old. Our estimates demonstrate that, for mothers without a university degree or with pre-reform earnings below the median, the program had effects on earnings and full-time employment even when the child was no longer eligible for Cash-for-Care at ages four and five. However, from age six, we can no longer see any effects. Further analysis suggests that the effects dissipate because most mothers remained attached to the labor force through part-time employment.

1. Introduction

Despite a dramatic increase in mothers' labor force participation since the 1960s, many mothers still choose to exit the labor market temporarily in order to stay at home with their young children. In the US, the labor market participation rate among mothers with their youngest child below age three is 54%, rising to 63% for mothers with their youngest child between ages three and five, and further to 73% among mothers with their youngest child above age five. In the OECD, the corresponding numbers for the three age groups are 51, 63 and 66%.[1]

In this paper, we investigate how mothers' decision to stay at home with young children affects their subsequent work careers. In particular, we address the concern that such temporary exits may lead to long-run deterioration of women's post-birth careers (Lalive and Zweimüller, 2009). This is important to countries considering policies that either encourage or discourage mothers to work while their children are young. The Scandinavian countries, for example, encourage female labor force participation by providing high quality publicly subsidized day care. In addition to allowing mothers to combine family life with work while the children are young, these family policies may have consequences for mothers' long-term labor force participation. High female labor force participation is considered important to maintain economic growth and sustainable pension systems (Burniaux et al., 2003). Moreover, female labor force participation secures family income and may hence prevent the detrimental effects on child outcomes of growing up in poverty (Yeung et al., 2002; Duncan et al., 2010; Dahl and Lockner, 2012).

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[1] Maternal employment by age of youngest child in 2008 (OECD, 2011).

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There are many reasons as to why a mother's decision to stay at home with young children could affect her long-term labor force participation. According to theories of human capital (Mincer and Polachek, 1974), a mother who exits the labor market while the children are young will accumulate less work-related human capital; the exit may even lead to skill depreciation.[2] Moreover, opportunities in the labor market are network dependent (Rees, 1966; Calvó-Armengol and Jackson, 2004). Presumably, a woman who is working while her children are young is more likely to develop a relevant network for her career than a mother staying at home with her children. Thus, when considering returning to work, a mother who temporarily exited the labor market might be less productive and have a more limited network than a mother who chose not to exit. This will likely be reflected in lower earnings and fewer job opportunities, which may discourage her from reentering the labor market. Mothers' reentry into the labor force could be additionally discouraged if staying at home with young children increases the accumulation of human capital related to home production (Becker, 1991).

Identifying a causal effect of a mother's temporary labor market exit while her children are young on subsequent career development is difficult because of omitted variable bias. In particular, women who choose not to work while their children are young may have lower career aspirations than mothers who choose to work, in ways that we cannot observe. It should not be surprising then if mothers who choose to stay at home while their children are young have less successful careers than mothers who choose to work. For identification, we study a unique, natural experiment in Norway that increased mothers' incentives to withdraw from the labor market in order to stay at home with their young children. The program, Cash-for-Care, was universal and paid any parent a significant allowance if they did not utilize a publicly subsidized child care slot for their one- or two-year-old child. The subsidy substantially decreased the labor force participation of mothers of one- and two-year-olds (Schøne, 2004a; Drange, 2012).[3] Our paper investigates how this labor market exit when the child was one and two years old affected the mothers' subsequent career development, after they were no longer eligible for the Cash-for-Care subsidy.

The analysis utilizes a comprehensive, longitudinal register database containing annual records for every person in Norway. We estimate difference-in-difference models that exploit differences in individuals' exposure to the program among families with similar structures and within similar birth cohorts. We find that the Cash-for-Care subsidy decreased full-time employment among mothers of two-year-olds by about four percentage points. Following the development in mothers' labor supply as their child grows older, we can see that, for mothers without a university degree or with pre-reform earnings below the median, there is still a significant reduction in full-time employment when the child is four years old and no longer eligible for the subsidy. In this sub-sample, the estimates suggest that about 60% of the mothers who exited full-time employment when their child was two years old and eligible for the subsidy have still not returned to full-time employment at age four, the year after subsidy eligibility expired. At age five, there is still a significant negative effect. However, at ages six and seven, we can no longer see any effect of the subsidy on mothers' full-time labor market attachment. Looking at earnings, we find similar results. For the sub-sample of mothers with high education or high earnings, there is no persisting effect of the Cash-for-Care subsidy at ages four and five.

Further analysis suggests that the effects of the Cash-for-Care subsidy dissipate because most mothers who exit full time employment while the children are young remain attached to the labor force through part-time employment. Norway is known for a very flexible labor market with a high share of part-time employment. Such part-time employment may limit the loss in human capital and relevant network when exiting full time employment.

This paper contributes to the literature studying the long-term effects of parental leave policies on women's labor market outcomes. Several studies show that job protection[4] associated with parental leave increases the likelihood that mothers return to the labor market when the children are older, and that it increases the job continuity with the pre-birth employer (Ruhm, 1998; Berger et al., 2004; Baker and Milligan, 2008). Notably, these studies study the effect of labor market exits while the children are younger than one year. Our paper is closely related to two recent contributions that investigate the effects of parental leave extensions when the child is older than one year. Lalive and Zweimüller (2009) take advantage of an increase in the duration of paid- and job-protected parental leave in Austria from one to two years. In a regression discontinuity analysis, they demonstrate that even if most mothers exhaust the full duration of their leave, there is no effect on their employment and earnings one year after the parental leave has been exhausted. Similar findings are reported in Schönberg and Ludsteck (2011), which studies five different German parental leave extensions. Utilizing a difference-in-difference approach, this study demonstrates that the extensions, in which the job protection period is as long as the maternity benefit period, have no effect on mothers' employment or earnings six years after childbirth. However, the one reform that extended the maternity benefit period from 10 to 22 months, without extending the job protection period, had a large negative effect on mothers' employment and earnings six years after childbirth.

The Cash-for-Care subsidy increased mothers' incentives to stay home with their children up to age three,[5] which extends beyond the two year job protection period after child birth in Norway. Consistent with Schönberg and Ludsteck (2011) our study suggests that an extension in paid parental leave that is not connected to an extension in job protection, may reduce maternal labor force participation even at ages when the child is no longer eligible for the program. In contrast to Schönberg and Ludsteck (2011), however, we find that this effect dissipates by ages six and seven. The paper proceeds as follows. Section 2 describes the Cash-for-Care program and gives an overview of the Norwegian institutional setting. Section 3 presents the empirical strategy, and Section 4 describes the data. The empirical results are reported in Section 5, and Section 6 concludes.

2. Institutional details

2.1. Norway's Cash-for-Care program

After the election for the Norwegian parliament in September 1997, an alliance consisting of the Christian Democratic party, the Centre party and the Liberal party formed a new government. One of the issues on their political agenda was the introduction of a Cash-for-Care subsidy. The Cash-for-Care Act was passed in the parliament in June 1998 (Norwegian Ministry for Children and the Family, 1998).

In August 1998, the government began awarding tax free cash allowances to parents who did not use publicly subsidized child care programs. Any family with a one- or two-year-old toddler could claim this allowance. The government stated that the main goals of the allowance

[2] Some restoration of human capital may take place if the mother chooses to reenter the workforce after a temporary exit. In US data, Mincer and Ofek (1982) find a relatively rapid growth in wages for women reentering the labor force, suggesting a restoration of previously eroded human capital.

[3] See also Naz (2004) who utilizes survey data and investigates how the subsidy affected specialization in families; Hardoy and Schøne (2008) who study the effect of the subsidy on marital stability; and Rønsen (2009) who compares labor force participation of mothers of one- and two-year-olds before and after the reform. Rønsen (2009) investigates effects of the Cash-for-Care on mothers with eligible children (one- and two-year-olds) several years *after* the subsidy was introduced, and refers to this as long term effects. We refer to long term effects as effects on mothers' labor supply when the treated child was no longer eligible for Cash-for-Care.

[4] The right to return to the same job after the parental leave is over.

[5] Norway's parental leave is generous, and allows parents to exit the labor force for one year following the birth of a child. Even if Cash-for-Care is not as generous as the parental leave, it extends significant benefits until a child's third birthday.

Child's age	1997	1998	1999	2000	2001	2002	2003	2004	2005
Age 1	0	0–5	0–11	0–11	0–11	0–11	0–11	0–11	0–11
Age 2	0	0–5	12	12	12	12	12	12	12
Age 3	0	0	1–12	1–12	1–12	1–12	1–12	1–12	1–12
Age 4	0	0	0	0	0	0	0	0	0
Age 5	0	0	0	0	0	0	0	0	0
Age 6	0	0	0	0	0	0	0	0	0
Age 7	0	0	0	0	0	0	0	0	0
Age 8	0	0	0	0	0	0	0	0	0
Age 9	0	0	0	0	0	0	0	0	0
Age 10	0	0	0	0	0	0	0	0	0

Notes: The number in each cell denotes months of Cash-for-Care eligibility. The darkly shaded cells represent fully treated children, whereas the lightly shaded cells represent partly treated children. The first cohort that is partly treated was born in 1996. The first cohort that is fully treated was born in 1998.

Fig. 1. Months of eligibility.

were to give families financial freedom to stay at home with their young children, to allow families themselves to choose what kind of care they wished for their children, and to equalize public transfers to families — regardless of what kind of care the family wanted or had access to for their child.

At the time of introduction, the Cash-for-Care allowance constituted a significant part of family earnings even for high-income families. The annual allowance was NOK 36,000,[6] and the average annual fee for publicly subsidized child care was about NOK 34,600 with some price subsidies for low-income families. Bettinger et al. (in press) demonstrate that for a family in the bottom income quartile, the effective after-tax price of a full-time day care slot for a one- or a two-year-old constituted about 40% of average family earnings. For the third and fourth income quartiles, the Cash-for-Care allowance constituted 15 and 10% respectively.

While the Cash-for-Care program was implemented simultaneously throughout Norway, there is variation in time and the ages of eligible children. Starting in August 1998, all one-year-old children were eligible for the Cash-for-Care allowance, from the month after they turned one year-old. From January 1999, both one- and two-year-old children were eligible.[7] As a consequence, all children born from 1998 onward were eligible for 24 months of the Cash-for-Care allowance. For these children, eligibility started at the end or close to the end of maternity leave. We will refer to these children as fully treated. Children born prior to 1996 are not affected by the Cash-for-Care allowance. Children born in 1996 or 1997 could be eligible for as little as one month and as many as 24 months of the Cash-for-Care allowance. We will refer to these children as partly treated.

Fig. 1 describes the nature of the treatment. Each cell represents the age of a child in a given year. Each cohort of children can be followed diagonally in this matrix. The darkly shaded cells represent fully treated children, whereas the lightly shaded cells represent partly treated children. The numbers in each cell denote how many months the mother of a child at a given age in a given year was eligible for the subsidy. Note that we have also shaded the cells of some of the older children not eligible for the Cash-for-Care subsidy. As we can see, these cells illustrate children who were treated as one or two-year-olds. If the Cash-for-Care subsidy had a persistent effect on the mothers' labor supply, we should see a treatment effect in these cells.

The uptake of the Cash-for-Care program was substantial. Fig. 2 shows the number of children who received the subsidy (the entire subsidy or parts of it) in the 1998–2001 period. As we can see, in 1999, the first year in which all families with one- and two-year-old children were eligible, three out of four families received a partial or full subsidy. It is also worth noting that the number of families receiving the subsidy is quite similar regardless of the age of the child. Only approximately 5000 families stopped receiving the subsidy when their child turned two years old.[8]

2.2. Norwegian parental leave, child care and female labor market participation

In the decade prior to the introduction of the Cash-for-Care subsidy, there was a substantial increase in female labor market participation in Norway. In 1991, 74% of married or cohabitant mothers with children aged 0–15 years were working. At the time of the introduction of the Cash-for-Care subsidy in 1998, this figure had increased to 81% (Kjeldstad and Rønsen, 2002). However, the incidence of part-time employment is above the OECD average.[9] Women are in general overrepresented in the public sector, particularly in occupations related to health and social work, where the prospect of obtaining a part-time job is good (Tronstad, 2007).

The Norwegian government introduced the Cash-for-Care program at a time of extensive use of publicly subsidized day care. About 40% of children aged one or two years used publicly subsidized day care,[10] and there was a short supply of these day care programs. Moreover, at the time of its introduction, parents were entitled to 42 weeks of parental leave with full compensation, or alternatively 52 weeks with 80% wage compensation,[11] in addition to one year of unpaid job protection for each parent in connection with child birth. The Cash-for-Care program made it less costly to extend the period at home with the child before returning to work. However, if a mother chose to stay home with her children until age three her job would not be protected.

If a family wanted to receive the Cash-for-Care allowance, they would either have to take care of their child themselves or use informal care (e.g., relative, neighbor, or home-based day care). In Norway, formalized care consists almost exclusively of public and publicly subsidized private day care centers. The two types of centers are regulated by the same law; they basically offer the same type of program, have the same price schedule for parental pay and are equally subsidized. As there were very few private day care centers that did not run publicly subsidized programs, Cash-for-Care recipients in practice did not have the option of using private formalized care.

The Cash-for-Care program gave families strong incentives to reduce labor supply and substitute formal care with parental care, or to substitute formal care with informal care. Rønsen (2001) demonstrates that the program increased both parental time at home and time in informal day care. Moreover, the Cash-for-Care allowance decreased eligible

[6] According to the National Bank of Norway 1 Euro = 8.1 NOK in 2000.

[7] There was an exemption from this rule for all children who turned two years old after August 1, 1998. This exemption ensured that no children had a break in eligibility from the Cash-for-Care allowance.

[8] We assume that families who received the subsidy for a two-year-old child also received it when the child was one.

[9] OECD Family Database, based on statistics from 2007 (OECD, 2011).

[10] OECD Labor Market Statistics: http://stats.oecd.org/ and Statistics Norway, Official Statistics of Norway: Kindergartens 1998.

[11] In 2009, parental leave was extended to 46 weeks of full compensation or 56 weeks of 80% compensation.

	All	One-year-olds	Two-year-olds	% of all children aged one or two
1998	61,243	47,983	13,260	50.1
1999	89,592	46,598	42,994	74.8
2000	88,234	46,988	41,243	74.3
2001	87,580	46,549	41,031	73.2

Notes: The subsidy was introduced for one-year-olds from August 1, 1998 and for two-year-olds from January 1, 1999. Source:Norwegian Welfare Administration

Fig. 2. Families receiving the Cash-for-Care subsidy.

mothers' full-time labor force participation by about four percentage points across the whole population but had no effect on fathers' labor force participation (Schøne, 2004a; Drange, 2012).

2.3. Other family reforms

During the years prior to the introduction of the Cash-for-Care allowance in 1998, Norway implemented several work–family-related policies. 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 and to 42 weeks in 1993. Moreover, in 1993, Norway introduced a paternity quota of paid parental leave. Of the 42 weeks of paid parental leave, four weeks were reserved exclusively for the father.[12]

Notably, the parental leave policies were initiated at least three years prior to the introduction of the Cash-for-Care allowance, and at least five years before the birth of the first cohort that was fully entitled to the allowance. As such, these policies should not be of any concern for our identification strategy if uptake was immediate. The uptake of the expansions in parental leave was immediate. However, the paternity quota was not extensively used until two years after implementation. The slow uptake of the paternity quota may raise a concern for our analyses because it implies that our partially treated younger siblings were fully affected by the paternity quota, whereas our control group was only partially affected by the quota. Empirical investigations of the paternity reform suggest, however, that the paternity quota had no significant effect on mothers' labor supply (Rege and Solli, in press; Cools et al., 2013).

Another reform relevant to our study is the 1997 school reform that changed the school starting age from age seven to age six. All children in our sample started school at age six, so even if mothers' labor supply might be affected by their children starting school earlier,[13] the same effect should be prevalent for both the comparison and treatment groups. The school reform may still be of concern because it led to an increase in the cover of publicly subsidized day care slots in 1997, because six-year-olds were no longer in need of child care slots. This increase in day care availability could possibly increase mothers' labor force participation while having young children[14] and thereby bias our estimates downward. However, considering the development in child care slots for the age groups in question, there seems to be little evidence of a spike in child care attendance for five-year-olds in 1997 or the following years, suggesting that child care for this age group was not rationed at the time.[15] For the two-year-olds, however, there is an increase in enrolled children in 1997. We carefully address this concern in the empirical strategy.

[12] Since 1993 the paternity quota has been extended multiple times. As the first extension was in 2005, these extensions did, however, not affect the cohorts in our study.

[13] Gelbach (2002) finds that mothers' labor supply increases when their oldest child starts school.

[14] See, for instance, Baker and Milligan (2008).

[15] Statistics Norway 2003 (http://www.ssb.no/nos_barnehager/nos_d328/tab/tab-3.html).

Finally, at the time of our study, the government implemented a reform giving single mothers stronger incentives to reenter the labor market. The reform was gradually implemented, starting January 1st 1998, and entailed work requirements and higher in-work benefit levels. According to Mogstad and Pronzato (2009), this led to an increase in single mothers' labor force attachment. Given the gradual phase-in of this policy change, it is hard to determine whether and how it would affect our treatment and control groups. We include a covariate capturing whether the family is a single-mother household in all analyses unless otherwise stated. Moreover, a robustness analysis, in which we exclude the group of single mothers, provides similar results as the main analysis. These analyses are not reported in the paper but are available from the authors upon request.

3. Empirical strategy

To estimate the effects of the Cash-for-Care program on mothers' long-term labor market outcomes, we exploit variation across similar families over time in a difference-in-difference analysis. The shading in Fig. 1 illustrates the nature of the treatment. We can see from Fig. 1 that children born in 1995 represent the latest cohort that is not treated, and children born in 1998 represent the first cohort that is fully treated. We estimate the following difference-in-difference coefficient for the effect of the Cash-for-Care subsidy when the children are a years old (for $a = 2, 3, 4, 5$ and 6 years old):

$$\gamma_a = \left(ls_{a,98+a} - ls_{a,95+a}\right) - \left(ls_{a+3,98+a} - ls_{a+3,95+a}\right) \tag{1}$$

where $ls_{a,y}$ denotes the labor force participation rate of mothers of children age a in year y. The difference-in-difference coefficient γ_a measures the changes in full-time attachment between year 1995 + a and year 1998 + a for mothers of a year-olds compared with mothers of a + 3 year-olds.

To understand better our difference-in-difference approach, let us start by considering how the introduction of the Cash-for-Care subsidy affected eligible mothers of two-year-olds. By substituting $a = 2$ in Eq. (1), we get:

$$\gamma_2 = \left(ls_{2,00} - ls_{2,97}\right) - \left(ls_{5,00} - ls_{5,97}\right). \tag{2}$$

The children from the 1995 and 1998 cohorts are two years old in 1997 and 2000, respectively. Thus, we can look at how Cash-for-Care eligibility affected mothers of two-year-olds by examining the difference in the labor supply of mothers of two-year-olds between 1997 and 2000, which is the first difference in Eq. (2). Of course, there are many factors aside from the Cash-for-Care subsidy that may affect mothers' labor force participation between 1997 and 2000. Note from Fig. 1 that in 2000, five-year-olds are the youngest cohort who never received the Cash-for-Care subsidy. Thus, we can control for these other factors by subtracting the difference in labor force participation of mothers of five-year-olds between 1997 and year 2000, which is the second difference in Eq. (2). If trends in labor force participation are identical for mothers of two and five-year-olds, then γ_2 will capture the causal effect of the Cash-for-Care subsidy on mothers of two-year-olds.

When investigating how the Cash-for-Care subsidy affected mothers of three-year-olds, we follow the same cohorts and look at the same difference-in-difference *one* year forward. This will give us the γ_3 difference-in-difference coefficient. Similarly, we obtain the γ_4 coefficient by following the same cohorts and look at the same difference in differences *two* years forward, and so on.

In our empirical analyses, we estimate γ_a by restricting the sample to mothers of two- and five-year-olds in 1997 and 2000, respectively, and then estimate the following difference-in-difference model:

$$ls_{a,i} = \alpha + \beta age_{a,i} + \lambda year_{98+a,i} + \gamma_a \left(age_{a,i} year_{98+a,i} \right) + \eta X_i + \varepsilon_i \qquad (3)$$

where $ls_{a,i}$ is mother i's labor market outcome when the child is a years old, $age_{a,i}$ is a dummy indicating whether the child is a years old, $year_{98+a,i}$ is a dummy indicating whether the year is 1998 + a, and X is a vector capturing a rich set of observable characteristics of the child, mother and father that may influence the mother's labor market outcome.[16] These variables are all observed prior to the introduction of the Cash-for-Care subsidy and are specified in Section 4.

The difference-in-difference model in Eq. (3) will provide an unbiased estimate of γ_a if the trends in labor supply of mothers of a year-olds and a+ 3 year-olds would have been the same in the absence of the Cash-for-Care subsidy. There are several ways in which this identifying assumption might be violated. Below we elaborate on how we address possible challenges to the identifying assumption.

3.1. Trends in labor force participation

We know that mothers increased their labor market participation during the 1990s. This increase was not necessarily the same for mothers of children of different ages. Moreover, we could imagine that changing labor market conditions affected mothers differently across time and across the age of the children. By thoroughly exploring pre-trends in labor force participation and by performing a placebo test, we investigate the validity of our identifying assumption. Notably, this investigation also addresses the possible concern of mothers' labor force participation being affected by diverging trends in child care availability due to the 1997 school reform that changed the school starting age from age seven to age six (see Section 2.3). Diverging trends in child care availability could also be a direct consequence of the Cash-for-Care subsidy. In particular, child care availability for five-year-olds could have been affected by the Cash-for-Care reform if the reform caused mothers of toddlers to not utilize child care, and this again affected the availability of child care for mothers of older children. If mothers of five-year-olds in 2000 worked more than mothers of five-year-olds in 1997 due to better accessibility of child care, then our difference-in-difference approach would produce a negative estimate even if work force participation among mothers of two-year-olds is not affected by the subsidy (mothers of the young children might send their child to informal care arrangements).[17] In a robustness test we have investigated this concern by including municipality level child care coverage rates for five-year-olds in 1997 and 2000 — separately and interacted with the child's age. The results are robust to including these additional covariates.[18]

3.2. Fertility

Policies implemented to mitigate costs connected to the birth of a new child might increase fertility.[19] A sample selection criterion for all four groups in this study (comparison and treatment group, before and after the policy change) is that the children do not have a new sibling prior to the year they turn seven years old.[20] A possible fertility effect of the subsidy could affect the composition of the groups differently through this sample selection criterion. In particular, the youngest cohorts have younger mothers when the policy is introduced, and are thus more likely to have new siblings if the reform affects fertility.[21] First, we address this concern by demonstrating that our results are robust to the inclusion of children with younger siblings. Second, we construct a sample of mothers with children born in the relevant years *without* imposing any selection criteria. We proceed to run the exact same regression as in the main specification, with the exact same covariates, but with the outcome being if the mother has a new baby by year seven. If the subsidy indeed increased fertility to a larger extent for the youngest cohorts, we should expect a positive coefficient when implementing this specification on the unrestricted sample.

3.3. Timing of fertility

If parents could anticipate the introduction of the subsidy and time the birth of their child, this could lead to selection into the treatment group and thereby bias the results. Keeping in mind from Section 3.1 that the final decision on the implementation of the subsidy came in June 1998, timing of birth in 1998 should not be of major concern.[22] There was, however, a public debate on the issue prior to the election in September 1997. One could imagine that some parents wanted to postpone conception to after the election when they possessed more information on whether the subsidy would actually be implemented. The result would be that some children that would otherwise have been born late 1997 were instead born in 1998. To explore possible fertility effects we will display the number of children born by month for the cohorts 1996 (benchmark), 1997 and 1998.

4. Data

We utilize registry data called FD-trygd, which is a combination of several Norwegian registry databases provided by Statistics Norway. Our dataset contains records for every Norwegian resident from 1992 to 2005. The data provides individual demographic information (marital status, spouse identifier, sex, age, number of children), socio-economic data (years of education, income, wealth), current employment status (full time, part time, minor part time, self-employed), indicators of participation in any of Norway's welfare programs and geographic identifiers for county, municipality and neighborhood of residence. Importantly, the data contain identifiers for mother and father, which allow us to match the children to their parents.

As described in our empirical analysis section, our main analytic sample consists of all children of two and five years old in 1997 and 2000, respectively. We further restrict the sample to children who have no younger siblings at age seven. We make this restriction in order to ensure that the older children in the control group do not have treated siblings.[23] Lastly, we restrict the sample to mothers who we can observe throughout all relevant periods, i.e. when the child is

[16] Alternatively, we could use a triple difference approach as in Schøne (2004a,b), analyzing the change in labor force participation for the same individual from a pre- to a post-period. This would give us somewhat larger estimates (results from triple difference approach are not reported; available from authors on request). The advantage of the double difference is that is allows us to control for a rich set of parental characteristics at baseline.

[17] Note that this would also be a reform effect, but the interpretation is different, and we should not expect any long term effects on labor supply for mothers of younger children.

[18] Results are available from the authors upon request.

[19] See for instance Lalive and Zweimüller (2009).

[20] We make this restriction in order to ensure that the older children in the control group do not have treated siblings.

[21] In fact, the cohort born in 1992 has already turned seven when the subsidy is introduced, so in this group selection due to the subsidy cannot happen by definition.

[22] All children in our sample are born in 1998 or prior to 1998.

[23] As discussed in Section 3, this restriction is potentially endogenous if the reform affected fertility. However, our empirical analyses address this concern by demonstrating that our results are robust to the inclusion of children with younger siblings.

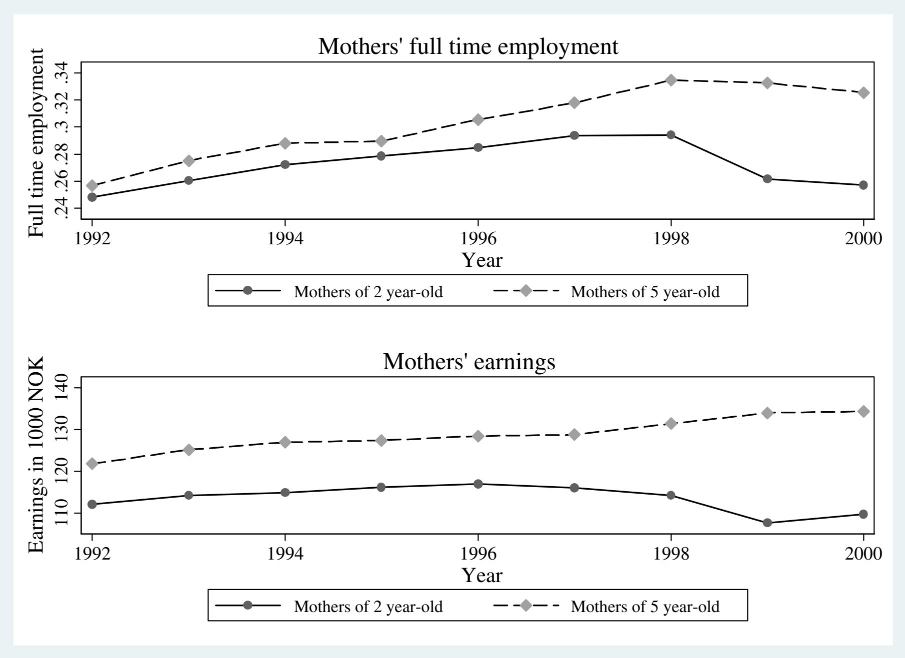

Fig. 3. Trends in mothers' full time employment and mothers' earnings.

two, three, four, five, six and seven years old (treatment group) and five, six, seven, eight, nine and ten years old (control group).

We analyze how the Cash-for-Care subsidy affected mothers' long-term labor market attachment by using several different outcome variables. All outcome variables are constructed for each of the years when the child is 2–7 years old for the children aged two years in 1997 and 2000, and for each of the years when the child is 5–10 years old for the children aged five years in 1997 and 2000. Our key outcome, which we refer to as *full-time employed*, is a variable denoting whether the mother is working more than 30 h per week at the end of the year. Inspection of the data reveals that some individuals are recorded as full-time employed despite very low recorded earnings or even zero earnings. This is likely because of lags in firms' submissions of employer information. We address this by coding everybody with earnings that precludes employment as non-employed.[24] We also use an employment variable capturing whether the mother was either full-time or part-time employed, which we refer to as *employed*. A mother is coded as employed if she is registered as working minor part time, part time or full time and has earnings above a certain threshold.[25] Finally, we use earnings as an outcome measure.[26] This variable includes all earnings that qualify for pension points in the Norwegian Social Insurance Scheme, i.e. income from employment and self-employment, as well as work related transfers such as sickness benefit, parental leave benefit, disability pension and unemployment benefit. We avoid using log earnings because all mothers are included in the earnings analyses, including those who are not working and have zero earnings. All mothers have to be included in the earnings analysis, as the sample of working mothers is endogenous to the Cash-for-Care reform.

Our data allows us to construct several control variables capturing important child, father and mother characteristics that we include in our regression analyses. In order to ensure that covariates are not endogenous to the reform, all covariates are collected from a baseline year, prior to the introduction of the Cash-for-Care subsidy. For the children aged two years in 1997 and 2000, we collect covariates from the year prior to their birth in 1994 and 1997. For the children aged five years in 1997 and 2000, we collect covariates when these children are two years old in 1994 and 1997. Since the Cash-for-Care subsidy was introduced in 1998, this assures that all covariates are measured prior to the introduction of the subsidy. Our control variables include the following: child gender, number of children (0, 1, 2, 3, 4, ≥5),[27] mother's age at birth of youngest child (<20, 20–24, 25–29, 30–34, 35–39, 40–44, ≥45), age at birth of oldest child (<20, 20–24, 25–29, 30–34, 35–39, 40–44, ≥45), father's age at birth of youngest child, parents' education (completed high school, completed college), linear and quadratic controls for parents' earnings, parents' employment status (minor part time, part time or full time), indicator for parents receiving any social welfare benefits, indicator for parents living in a densely populated area (city), indicator for parents' immigration status, indicator for being a single mother, and municipality-specific unemployment rates interacted with the age of the children. Finally, we include municipality-fixed effects.

5. Empirical results

5.1. Summary statistics

Fig. 3 shows the trends in mothers' full-time employment[28] and mothers' earnings throughout the 1990s. Included are mothers of two-

[24] Everybody with an income less than 2 G is coded as not being full-time employed. G is set by the Norwegian Labour and Welfare Administration every year and is included in most formulae for welfare transfers. In 1997, G was NOK 42,000.

[25] Earnings above 0.25 G.

[26] Earnings for mothers are inflation adjusted to the 1997 level by the change in earnings in the entire female population (aged 20–67).

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

[28] Because of a change in the registration procedure in Statistics Norway, there is a surge in missing values for the labor supply variable in 1998 and 1999. The share of missing values is similar for mothers of children of different ages and is, according to Statistics Norway, a result of a change in registration routines during these particular years. To ensure that our results are not biased by these changes, we run regressions with various definitions of full-time employment based on earnings. These analyses provide similar results and are available from the authors upon request.

Table 1
Summary statistics.
Source: Administrative registers: FD Trygd.

	Treat/pre	Treat/post	D	Comp./pre	Comp/post	D	D-in-D
	Born 1995	Born 1998		Born 1992	Born 1995		
Panel A: Outcome var.							
Mother full-time age 2/5	0.300	0.264	−0.036**	0.328	0.335	0.007	−0.044**
Mother full-time age 4/7	0.333	0.369	0.036**	0.353	0.413	0.060**	−0.024**
Mother full-time age 7/10	0.412	0.417	0.005	0.436	0.448	0.012**	−0.007
Mother earnings age 2/5	117,100	110,826	−6274**	130,864	136,271	5407**	−11,681**
Mother earnings age 4/7	131,989	134,834	2845**	140,527	146,629	6102**	−3257**
Mother earnings age 7/10	146,458	147,995	1537*	152,499	154,941	2442**	−905
Panel B: Child characteristics							
2 children	0.452	0.450	−0.002	0.441	0.452	0.010**	−0.012*
3 children	0.266	0.269	0.003	0.263	0.266	0.003	0.000
4 children	0.063	0.064	0.001	0.066	0.062	−0.004	0.005+
5 children or more	0.020	0.020	0.000	0.020	0.019	−0.001	0.001
Gender = girl	0.491	0.493	0.002	0.488	0.491	0.004	−0.002
Panel C: Parent characteristics							
M prior earnings	139,004	139,624	620	115,718	116,814	1096	−477
M minor part-time prior	0.122	0.129	0.007*	0.156	0.165	0.008**	−0.002
M part-time prior	0.113	0.118	0.005+	0.125	0.132	0.007**	−0.002
M full-time prior	0.401	0.427	0.026**	0.294	0.323	0.029**	−0.003
M high school	0.513	0.580	0.066**	0.465	0.529	0.064**	0.002
M college	0.244	0.283	0.038**	0.231	0.255	0.024**	0.015**
M age	30.506	31.012	0.507**	30.102	30.505	0.403**	0.104+
M immigrant	0.068	0.075	0.007**	0.065	0.073	0.008**	−0.001
M on welfare	0.050	0.049	−0.001	0.062	0.061	−0.002	0.001
M urban area	0.757	0.755	−0.002	0.749	0.761	0.012**	−0.014**
Single mum	0.148	0.147	−0.001	0.165	0.148	−0.017**	0.016**
Unemployed	0.028	0.023	−0.005**	0.028	0.023	−0.005**	0.000
F prior earnings	247,514	249,342	1828	266799	262,752	−4047**	5875**
F minor part-time prior	0.023	0.029	0.006**	0.019	0.021	0.003*	0.003+
F part-time prior	0.019	0.019	−0.001	0.016	0.018	0.002	−0.002
F full-time prior	0.691	0.728	0.038**	0.700	0.732	0.031**	0.006
F high school	0.553	0.608	0.055**	0.543	0.583	0.040**	0.015**
F college	0.226	0.248	0.022**	0.235	0.237	0.001	0.021**
F age	33.409	33.830	0.421**	33.020	33.409	0.388**	0.032
F immigrant	0.072	0.075	0.003	0.070	0.076	0.006**	−0.003
F on welfare	0.056	0.047	−0.009**	0.064	0.060	−0.005**	−0.004
F urban area	0.747	0.747	0.000	0.737	0.748	0.011**	−0.011*
N	29,640	29,365		28,533	29,642		

Notes: Mean or share of indicated variable with differences. Earnings are inflation adjusted with 1997 as base year (in NOK). +, * and ** denote significance at the 10, 5 and 1% levels respectively (two-sided *t*-test).

and five-year-olds in a given year. The first cohort of partly treated children turns two in 1998, while the first cohort of fully treated children turns two in 2000. Looking at the trends in full-time employment, we first note that, consistent with our identifying assumption, the trends seem fairly similar for the years prior to the policy change in 1998. Then there is a relative drop in the labor force participation of mothers of two-year-olds compared with mothers of five-year-olds in 1998–2000. This is consistent with an effect of the Cash-for-Care subsidy on mothers' full-time employment when the child is two years old. Looking at the trends in earnings, we see a similar pattern. The pre-reform trends are similar for the two age groups. For mothers of two-year-olds, there is a relative drop in earnings in 1998 and 1999, compared with mothers of five-year-olds.

Table 1 provides summary statistics for full-time employment and earnings at different child ages, and for key background characteristics.[29] First, focusing on the outcome variables in Panel A, we can see a decrease in full-time labor force participation among mothers of two-year-olds between post- and pre-treatment, whereas mothers of five-year-olds slightly increased their full-time labor force participation between post- and pre-treatment. The unadjusted difference-in-difference estimate is significant and consistent with an effect of the Cash-for-Care subsidy on mothers' full-time employment when the child is two years old. There is also some evidence consistent with persisting effects of the Cash-for-Care subsidy at age four, after subsidy eligibility has expired. We can see that mothers of four-year-olds increased their full-time employment between post-reform and pre-reform, but less so than mothers of seven-years-olds. The unadjusted difference-in-difference estimate is significant. However, this effect seems to have dissipated when the treated children reached age seven, when the unadjusted difference-in-difference estimate is smaller and no longer statistically significant. Not surprisingly, looking at differences in earnings we see the same pattern, reflecting the changes in full time employment for the different groups of mothers.

When considering the covariates in Table 1, Panels B and C, there seem to be a few changes among parents of treated children that do not correspond to changes among the parents of children in the comparison group. There seems to be an increase in the number of families with two children in the comparison group post-reform. Furthermore, we see that mothers and fathers with children from later cohorts are more likely to have finished college than are mothers and fathers of the older children. This possibly relates to general trends in education of the population during the period.[30] We also note that while the share of single mothers is stable in the treatment group, it decreases somewhat in the comparison group.

[29] See Section 4 for a detailed list of all covariates included in the analysis.

[30] Statistikkbanken, Statistics Norway.

Table 2
The effect of Cash-for-Care on mothers' full-time employment.
Source: Administrative registers: FD Trygd.

	Model 1	Model 2	Model 3	Model 4	Model 5	Model 6
Age 2	−0.041**	−0.040**	−0.040**	−0.040**	−0.040**	−0.034**
	(0.006)	(0.006)	(0.004)	(0.004)	(0.004)	(0.005)
Age 3	−0.039**	−0.038**	−0.038**	−0.037**	−0.037**	−0.033**
	(0.006)	(0.006)	(0.005)	(0.005)	(0.005)	(0.005)
Age 4	−0.022**	−0.021**	−0.021**	−0.020**	−0.020**	−0.017**
	(0.006)	(0.006)	(0.005)	(0.005)	(0.005)	(0.005)
Age 5	−0.010	−0.009	−0.009+	−0.009+	−0.009+	−0.007
	(0.006)	(0.006)	(0.005)	(0.005)	(0.005)	(0.006)
Age 6	−0.007	−0.006	−0.006	−0.006	−0.006	−0.003
	(0.006)	(0.006)	(0.006)	(0.006)	(0.006)	(0.006)
Age 7	−0.005	−0.004	−0.005	−0.004	−0.004	−0.000
	(0.006)	(0.006)	(0.006)	(0.006)	(0.006)	(0.006)
N = 117,180						
Included covariates						
Child char.	X		X	X	X	X
Mother char.			X	X	X	X
Father char.				X	X	X
Unemployment					X	X
Unemployment × age						X

Notes: +, * and ** denote significance at the 10, 5 and 1% levels respectively. Estimates are based on OLS on Eq. (2) with the dependent variable being whether the mother is full time employed. We estimate the model separately for different ages. We follow the cohort of fully treated children born in 1998 as two-year-olds in 2000, three-year-olds in 2001, etc. For all ages, Model 1 is run without covariates. In the following five models, we add the following variables: child's characteristics, mother's characteristics, father's characteristics, municipality-specific unemployment rate in the year when the treated children turned two, and in Model 6, the same unemployment rate interacted with the age of the child. Robust standard errors (in parenthesis) are clustered on the child's mother and account for heteroscedasticity and nonindependence of residuals across mothers' labor force participation observed at different points in time. All specifications include municipality fixed effects.

Finally, there is a barely significant decrease in the share of parents who lived in densely populated areas. This might influence the labor force attachment if there is a higher unemployment rate in rural areas.

As described in our empirical strategy, we account for possible observable changes in the composition of the post- and pre-reform groups by including a rich set of parental and child characteristics (described in Section 4). We also add covariates sequentially in order to investigate if differential trends or observed changes in the composition of the groups affect our estimates. Moreover, we address the concern of differential effects of different labor market conditions by including controls for the local unemployment rate interacted with the age of the child.

5.2. Mothers' labor supply

Table 2 presents our main results. In Model 1, we report the unadjusted difference-in-difference estimate for how the Cash-for-Care reform affected full-time employment of mothers of fully treated children (born 1998) at different ages for the child. We can see a large and significant effect of the Cash-for-Care reform on mothers' labor force participation in the years when the child turns two and three years old. This is the age at which the parents are fully or partly eligible for the subsidy. The estimate suggests that the subsidy decreased full-time employment among mothers of two-year-olds by about four percentage points, which is similar to the estimates in other studies (Schøne, 2004a; Drange, 2012).[31]

Following the development in mothers' labor supply as their child grows older, we observe that when the affected child is four years old and no longer eligible for the subsidy, there is still a significant effect on mothers' full-time employment. The subsidy decreased full-time employment among mothers of four-year-olds by about two percentage points. This suggests that about 50% of the mothers who exited full-time employment when their child was two years old and eligible for the subsidy were still not working full time in the year after the subsidy expired. When the treated child is five–seven years old, we can see that the difference-in-difference estimates are smaller than at age four and no longer significant.

As discussed in the Empirical strategy section, one concern for the validity of our identifying assumption could be compositional changes among the different groups of mothers. In Models 2, 3, and 4, we investigate robustness by stepwise adding covariates capturing the child and parental characteristics described in Section 4. We can see that adding child-, mother- and father-specific characteristics has very small impacts on the estimates. This indicates that our estimates are not biased by compositional changes among the different groups of mothers.

Another concern for our identifying assumption is unemployment shocks that affect mothers differently depending on the age of their children. We investigate this concern in Models 5 and 6 by adding controls for local unemployment rates, linearly and interacted with the child's age. We can see that the effect estimates are robust to these inclusions, suggesting that the estimates are not biased by local unemployment shocks that affect mothers with children of different ages differently.

5.3. Specification analysis

In the specification analysis in Table 3, we further investigate the validity of our identifying assumption. The matrix reports unadjusted difference-in-difference estimates of effects on labor supply for mothers with children of different ages across different years. The comparison year is 1996 and the comparison group is mothers of eight-year-olds (mothers of eight-year-olds were not affected by the subsidy during the years included in the matrix): The matrix compares differences in labor supply between mothers of eight year-olds and mothers of children aged two to seven; and how these differences change across years. To understand the matrix better, consider the estimate at age two in year 2000 of negative 0.048. This difference-in-difference estimate corresponds to the first estimate in Model 1 in Table 2 with two important differences: in the matrix, the comparison year is 1996 instead of 1997, and the comparison group is mothers of eight-year-old children instead of mothers of five-year-olds.[32]

This matrix has two purposes. First, it demonstrates that our estimates are robust to other choices of comparison year and age. Second, the matrix allows us to investigate carefully whether the treatment effect appears in a way that is consistent with the introduction of the Cash-for-Care reform. In particular, if our effect estimates in Table 2 are due to the Cash-for-Care reform, then we should see a pattern in Table 3 that is similar to the shading in Fig. 2. We should not see any effects in the non-shaded cells, as these are mothers of non-treated children. We may see some effects in the lightly shaded cells, as these are mothers of partly treated children. We should see effects in the darkly shaded cells, at least for mothers of two- and three-year-olds, as these are mothers of fully treated children. If the reform has persisting effects, then we should also see an effect on older children in darkly shaded cells.

We can see that the estimates appear largely consistent with the introduction of the Cash-for-Care reform. All the coefficients in the darkly shaded cells with fully treated children are large and significant, even at ages four and five. We can also see that several coefficients in the lightly

[31] Notably, at age three the effect of the Cash-for-Care is virtually the same as for age two. This may seem odd since the age three children are in average only eligible for half a year. In Tables 5 and 6 we can see that with earnings or "any labor market attachment" as dependent variables, the treatment effects are substantially larger at age two than at age three.

[32] Additionally, the robustness matrix has a different sample selection criterion. In order to include all ages of children across a large number of years, we include all children with no younger siblings in the outcome year (Table 2 uses children with no younger siblings at age seven).

Table 3
Specification analysis.
Source: Administrative registers: FD Trygd.

	1996	1997	1998	1999	2000	2001	2002	2003
Age 2	–	−0.010+	−0.020**	−0.042**	−0.048**	−0.059**	−0.060**	−0.055**
		(0.005)	(0.005)	(0.005)	(0.005)	(0.005)	(0.005)	(0.005)
Age 3	–	−0.005	−0.002	−0.017**	−0.029**	−0.036**	−0.044**	−0.035**
		(0.005)	(0.005)	(0.005)	(0.005)	(0.006)	(0.005)	(0.005)
Age 4	–	−0.003	0.001	−0.002	−0.011+	−0.018**	−0.028**	−0.024**
		(0.005)	(0.005)	(0.005)	(0.006)	(0.005)	(0.005)	(0.005)
Age 5	–	−0.009+	−0.004	0.001	−0.003	−0.007	−0.020**	−0.017**
		(0.006)	(0.006)	(0.006)	(0.005)	(0.006)	(0.006)	(0.006)
Age 6	–	−0.002	−0.000	0.001	0.003	0.005	−0.005	−0.007
		(0.006)	(0.006)	(0.006)	(0.006)	(0.006)	(0.006)	(0.006)
Age 7	–	−0.007	−0.001	−0.003	−0.003	0.002	−0.007	−0.002
		(0.007)	(0.006)	(0.006)	(0.006)	(0.006)	(0.006)	(0.006)
Age 8	–		–	–	–	–	–	–
N		1,999,083						
R sq		0.003						
Mean		0.359						

Notes: +, * and ** denote significance at the 10, 5 and 1% levels respectively. Estimations are based on one OLS estimation. The matrix reports the unadjusted difference-in-difference estimates at different ages for the child in different years Changes in all years are measured relative to 1996, and the reference age is eight. Because of data constraints, this matrix is run without covariates. A child will be excluded from the sample in the year when the mother gives birth to a new sibling. Standard errors (in parentheses) are clustered on the child's mother and account for heteroscedasticity and nonindependence of residuals across mothers' labor force participation observed at different points in time.

shaded cells are significant, in particular for the cohort of children born in 1997, among whom many were close to being fully treated. Important for our identifying assumption, we can see that in most of the cells with no shading, the difference-in-difference estimates are small and insignificant. This suggests that trends in labor supply between mothers of children of different ages are similar prior to the reform, supporting our identifying assumption. Notably, there is a weakly significant difference-in-difference estimate at age two in 1997. This may reflect behavioral changes in expectation of the Cash-for-Care subsidy.

5.4. Subsample analysis and alternative dependent variables

In Table 4 we investigate the differential effects of Cash-for-Care across different levels of education and pre-birth earnings. All subsample analyses use our preferred model from Table 2, Model 6. Several interesting patterns emerge. Comparing Models 1 and 2, we can see that Cash-for-Care subsidy affects the full-time employment of mothers of two- and three-year-olds similarly across education level.[33] Moreover, comparing Models 3 and 4 we can see that subsidy affects mothers of two- and three-year-old children similarly across baseline earnings level. Even if the short term effects are similar, we observe that when the child is four and five years old the negative effects of the Cash-for-Care subsidy are only prevalent among mothers without college education and mothers who were low earners at baseline. Mothers with college education and mothers with above median earnings at baseline are back in a full-time position to the same extent as the non-treated mothers.

In Table 5, we investigate the effect of the Cash-for-Care subsidy on mothers' earnings. The structure of this table is identical to Table 4; however, full-time employment is replaced with earnings as the dependent variable. Importantly, because employment is endogenous to the reform, we include the same sample of mothers and do not focus on employed mothers only. Interestingly, we see in Model 1 that there are no effects of the Cash-for-Care subsidy on earnings for highly educated mothers when they are no longer eligible for the subsidy. This is somewhat surprising, as above 3% of these mothers exited full-time employment when eligible for the Cash-for-Care subsidy (Model 1, Table 4). As discussed in the Introduction, we know from theories of human capital that career breaks or part time employment may lead to a relative decrease in work-related human capital accumulation, which should be reflected in subsequent earnings.

In Model 2, we see significant effects of the subsidy on earnings for less-educated mothers of four year-old children. For mothers of four-year-olds, the estimate corresponds to a reduction of about 2% compared to the mean value of earnings in the group of mothers of four year-olds in 1997 (NOK 113,287). This reduction in earnings may reflect a relative loss in human capital in addition to the reduction in full time employment that we can see in Model 2 in Table 5. In Models 3 and 4, we can see a similar result. For the mothers with baseline earnings above the median, there are no effects of the Cash-for-Care subsidy on earnings at ages four and five. However, for mothers with earnings below the median, there seems to be indication of a negative estimate at age four, but the standard errors are large, and we cannot rule out a zero effect.

Table 4
Subsample analysis: Full-time employment.
Source: Administrative registers: FD Trygd.

	Model 1	Model 2	Model 3	Model 4
	Mother finished college	Mother not finished college	Former earnings below median	Former earnings above median
Age 2	−0.033**	−0.033**	−0.025**	−0.034**
	(0.010)	(0.005)	(0.005)	(0.008)
Age 3	−0.034**	−0.034**	−0.037**	−0.027**
	(0.011)	(0.006)	(0.006)	(0.008)
Age 4	−0.006	−0.022**	−0.027**	−0.008
	(0.011)	(0.006)	(0.007)	(0.008)
Age 5	0.004	−0.012*	−0.014*	−0.008
	(0.012)	(0.006)	(0.007)	(0.008)
Age 6	0.010	−0.008	−0.009	−0.003
	(0.012)	(0.007)	(0.007)	(0.008)
Age 7	0.011	−0.006	−0.002	−0.005
	(0.012)	(0.007)	(0.008)	(0.009)
N	29,682	87,498	58,604	58,576

Notes: +, * and ** denote significance at the 10, 5 and 1% levels respectively. Estimates are based on OLS of Eq. (2) with the outcome being whether the mother has a full-time attachment to the labor market. We estimate the model separately for different ages. We follow the cohort of fully treated children born in 1998 as two-year-olds in 2000, three-year-olds in 2001, etc. The covariates described in Section 4 are included. Robust standard errors (in parentheses) are clustered on the child's mother and account for heteroscedasticity and nonindependence of the residuals across mothers' labor force participation observed at different points in time. All regressions include municipality fixed effects.

[33] Note, however, that even though the effects are similar at the margin, the percentage reduction in each subsample is different. The mean of the full-time employed among the college educated is 0.45. Thus, the reduction in this group is 8%, while in the group without a college education, the reduction is 15% (mean is 0.25).

Table 5
Subsample analysis: Earnings.
Source: Administrative registers: FD Trygd.

	Model 1	Model 2	Model 3	Model 4
	Mother finished college	Mother not finished college	Former earnings below median	Former earnings above median
Age 2	−6283**	−9133**	−8586**	−6994**
	(1587)	(808)	(989)	(1069)
Age 3	−3124+	−4916**	−5674**	−2530*
	(1764)	(920)	(1119)	(1181)
Age 4	−463	−1948*	−2244+	127
	(1834)	(955)	(1186)	(1193)
Age 5	748	−576	−224	279
	(1925)	(997)	(1231)	(1251)
Age 6	1999	158	652	749
	(2014)	(1034)	(1270)	(1306)
Age 7	2378	−410	393	−592
	(2001)	(1007)	(1251)	(1269)
N	29,682	87,498	58,604	58,576

Notes: +, * and ** denote significance at the 10, 5 and 1% levels respectively. Estimates are based on OLS of Eq. (2) with the outcome being mothers' linear earnings. Earnings are inflation adjusted with 1997 as the base year (in NOK), and censored at the 99th percentile. We estimate the model separately for different ages. We follow the cohort of fully treated children born in 1998 as two-year-olds in 2000, three-year-olds in 2001, etc. The covariates described in Section 4 are included. Robust standard errors (in parentheses) are clustered on the child's mother and account for heteroscedasticity and non-independence of residuals across mothers' labor force participation observed at different points in time. All regressions include municipality fixed effects. 1 Euro = 8.1 NOK (2000).

Comparing to the results in Table 4, Table 5 suggests that mothers who were closely attached to the labor market prior to birth and who chose to exit at ages one and two did not experience a reduction in human capital. The reduction in earnings for mothers who had a weaker attachment to the labor market may reflect a relative loss in human capital in addition to the reduction in full-time employment. To understand the differences in long term effects across subsamples better, we now turn to Table 6, where we investigate the effect of the Cash-for-Care subsidy on mothers' employment. This table has the same structure as Tables 4 and 5; however, the dependent variable is replaced with employment, which includes minor part-time and part-time in addition to full-time employment.

Table 6
Subsample analysis: Employment.
Source: Administrative registers: FD Trygd.

	Model 1	Model 2	Model 3	Model 4
	Mother finished college	Mother not finished college	Former earnings below median	Former earnings above median
Age 2	−0.009	−0.040**	−0.047**	−0.012**
	(0.006)	(0.005)	(0.007)	(0.003)
Age 3	−0.001	−0.008	−0.011	0.002
	(0.006)	(0.005)	(0.007)	(0.003)
Age 4	0.003	−0.004	−0.001	0.003
	(0.005)	(0.005)	(0.007)	(0.003)
Age 5	0.004	0.003	0.006	0.006*
	(0.005)	(0.005)	(0.007)	(0.003)
Age 6	0.003	0.004	0.005	0.008*
	(0.005)	(0.005)	(0.006)	(0.003)
Age 7	−0.005	0.004	0.005	−0.002
	(0.005)	(0.004)	(0.006)	(0.003)
N	29,682	87,498	58,604	58,576

Notes: +, * and ** denote significance at the 10, 5 and 1% levels respectively. Estimates are based on OLS of Eq. (2) with the dependent variable being whether the mother works full time, part time or minor part time. We estimate the model separately for different ages. We follow the cohort of fully treated children born in 1998 as two-year-olds in 2000, three-year-olds in 2001, etc. The covariates described in Section 4 are included. Robust standard errors (in parentheses) are clustered on the child's mother and account for heteroscedasticity and nonindependence of residuals across mothers' labor force participation observed at different points in time. All regressions include municipality fixed effects.

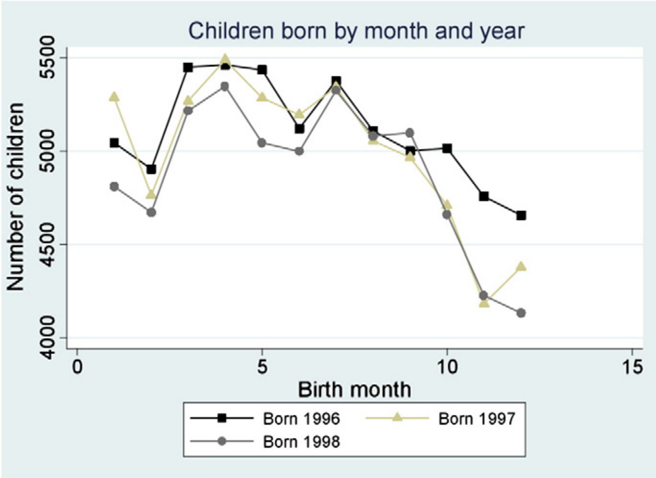

Fig. 4. Children born by year and month.

We can see that Table 6 displays a quite different picture compared to Tables 4 and 5. The Cash-for-Care subsidy has very small effects on employment already by age three in all sub-groups. This suggests that most mothers remained attached to the labor market through part-time employment, even if they withdrew from full time employment for an extended period. This finding is consistent with the findings in Schøne (2004b) who investigates how the Cash-for-Care subsidy affected employment of mothers with children age one-three. In Models 1 and 4 we can see that this is particularly true for mothers with college education and high earning mothers. This part time attachment to the labor force may provide one explanation for why these mothers did not experience a persistent decrease in earnings (see Table 5), despite the fact that many exited full time employment when eligible for the Cash-for-Care (see Table 4).

Table 7
Robustness.
Source: Administrative registers: FD Trygd.

	Model 1	Model 2	Model 3
	Full time employment	Full time employment	Having a new child by age 7
Age 2	−0.009	−0.029**	
	(0.006)	(0.003)	
Age 3	0.002	−0.024**	
	(0.006)	(0.003)	
Age 4	−0.006	−0.016**	
	(0.006)	(0.003)	
Age 5	0.001	−0.008*	
	(0.006)	(0.004)	
Age 6	0.005	0.009*	
	(0.006)	(0.004)	
Age 7	−0.001	0.011**	0.001
	(0.006)	(0.004)	(0.005)
N	118,034	239,698	239,698
Sample	Untreated cohorts born 1989, 1992 and 1995	Same as main model, but without restrictions	Same as main model, but without restrictions

Notes: +, * and ** denote significance at the 10, 5 and 1% level. Model 1 reports results from estimations based on OLS on Eq. (2) with outcome being full time employment and the sample is based on the same selection criteria as the main sample, but belonging to non-treated cohorts (born 1989, 1992 and 1995). Due to data restrictions, Model 1 does not include covariates or municipality fixed effects. Model 2 and 3 are estimations on a full sample including the entire cohorts of children, and include the covariates listed in Table 1. Model 2 reports results from estimations based on OLS on Eq. (2) with outcome being full time employment, whereas Model 3 report results from estimations based on OLS on Eq. (2) with outcome being if the mother has a new child by age 7. In all models we estimate the model separately for different ages. Robust standard errors (in parenthesis) are clustered on the child's mother and account for heteroscedasticity and non-independence of residuals across parents' labor force participation observed at different points in time. Model 2 and 3 include municipality fixed effects.

Table 8
Robustness: Covariates as outcome variables.
Source: Administrative registers: FD Trygd.

Model 1	Model 2	Model 3	Model 4	Model 5	Model 6
Mother full time baseline	Father full time baseline	Mother income baseline	Father income baseline	Mother finished high school	Father finished high school
0.002	0.003	−1011	4844**	−0.000	0.007
(0.005)	(0.005)	(848)	(1484)	(0.006)	(0.006)
N = 117,180.					

Notes: +, * and ** denote significance at the 10, 5 and 1% levels respectively. Estimates are based on OLS of Eq. (2) with the outcome being a new covariate in each model, respectively: Mother and father former full time employment, mother and father former earnings and mother and father high school completion share. Other covariates described in Section 4 are included. Robust standard errors (in parentheses) are clustered on the child's mother (father) and account for heteroscedasticity and nonindependence of the residuals across mothers' (fathers') labor force participation observed at different points in time. All regressions include municipality fixed effects.

In Models 2 and 3 we can see that also mothers with low education and mothers with low earnings remain attached to the labor force through part-time employment, but less so than high educated and high earning mothers. This may provide an explanation for why we see a more persistent effect of the Cash-for-Care subsidy on these mothers labor supply (see Tables 4 and 5).

5.5. Further robustness

We start by considering possible selection into the treatment group, i.e. into being born 1998. Parents that would otherwise have wanted to conceive a child early 1997 might have waited until after the election in order to have greater certainty about the subsidy. This would result in a reduction of children born late 1997, and a surge in children born second half of 1998. Fig. 4 below displays the number of children born by month in the years 1996, 1997 and 1998. Little indicates that the reform affected timing of birth; there are small variations throughout the year, but nothing stands out.

To further explore our identifying assumption, the first model in Table 7 provides estimates from a placebo model in which we assume that the reform took place three years earlier. The estimated model is identical to our preferred model 6 in Table 2 with one important difference; all observations and sample restrictions are moved three years forward in time. If preexisting trends affected labor supply differently in the treatment and comparison groups, we would discern a treatment effect in the years before the Cash-for-Care subsidy was introduced. It is clear from the first column in Table 7 that there is no indication of a preexisting diverging trend. This holds as the children grow older as well.

Another concern raised in Section 3 was that the sample selection criteria could bias the results. We are in particular worried about a possible subsidy effect on fertility. As discussed in Section 3, we restrict the sample to children who have no younger siblings at age seven in order to ensure that the older children in the control group do not have treated siblings, which may bias our estimates downward. This restriction may raise an endogeneity concern if the Cash-for-Care reform affected fertility. In Model 2 in Table 7, we demonstrate that our results are robust to the inclusion of children with younger siblings. As expected, the estimates are somewhat smaller, but we still see negative and significant effects on full-time employment among mothers of children 2–5 years old.

Moreover, in Model 3 in Table 7 we report results from a regression on whether the child has a new sibling by year seven. It is reassuring that the coefficient is a precisely estimated zero, suggesting that the cohorts included in treatment and comparison groups are not experiencing different fertility trends by year seven due to the subsidy.

As mentioned when discussing Table 2, a concern in difference-in-difference models is changes in the composition of included groups over time. An alternative approach to address this concern is to estimate our preferred model (Model 6 in Table 2); however, using key covariates at baseline as the dependent variable, and – of course – omitting these variables as covariates. Table 8 reports the results from this specification analysis, utilizing the following baseline variables as the dependent variable for both the mother and the father: Full time employment, earnings, and indicator for high school completion. We can see that none of the estimates are significant, except for the estimate on fathers' baseline earnings (Model 4 in Table 8). From Table 2 we know that fathers of children in the comparison group post reform have lower prior earnings. We rely on that controlling for fathers' prior earnings will address the sample composition issue.

6. Conclusion

In this paper we address the concern that temporary labor market exits while the children are young may lead to long-run deterioration of women's post-birth careers. We investigate how mothers' decision to stay at home with young children affects their subsequent work careers. For identification, we utilize the introduction of a Cash-for-Care subsidy, which was universal and paid any parent a significant allowance if they did not utilize publicly subsidized child care for their one- or two-year-old child. The subsidy substantially decreased the labor force participation of mothers of one- and two-year-olds (Schøne, 2004a,b; Drange, 2012). We demonstrate that, for mothers without a university degree or with pre-reform earnings below the median, the program had effects on earnings and full-time employment even when the child was no longer eligible for Cash-for-Care at ages four and five. However, from age six, we can no longer see any effects. For the sub-sample of mothers with high education or high earnings, there is no effect of the Cash-for-Care subsidy at ages four and five.

Further analysis suggests that the effects of the Cash-for-Care subsidy dissipate because most mothers who exit full time employment while the children are young remain attached to the labor force through part-time employment. Norway is known for a very flexible labor market with a high share of part-time employment. Such part-time employment may limit the loss in human capital and relevant network when exiting full time employment. Our findings are consistent with several studies that have emphasized the importance of a continued attachment to the labor market in securing labor supply post-birth.[34]

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[34] See, for instance, Berger et al. (2004), Baker and Milligan (2008) and Schönberg and Ludsteck (2011).

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