Universal Childcare for the Youngest and the Maternal Labour Supply

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DISCUSSION PAPER

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Abstract:

In this paper, we investigate whether the expansion of childcare leads to an increase in the female labour supply. We measure female labour supply at both the extensive and intensive margin. For identification, we exploit a nationwide reform that expanded childcare for 1–2-year-olds in Norway. Our results reveal a significant increase in the overall employment of mothers in the target group, but only weak evidence of an increase in contracted hours of work. However, both adjustments are only short term following the reform. When we consider sub-groups of mothers more closely, we find substantial heterogeneity in the affected outcomes and the timing of these effects. In particular, when we exclude mothers on job-protected maternity leave and with currently zero hours of work from the target group, we estimate even larger effects on employment and now significant effects on actual hours of work. For mothers with more than one child, we find significant long-term effects of the reform on both employment and hours of work.

Keywords: childcare, female labour supply, contracted hours, actual hours, causal effects.

JEL Codes: J08, J13, J22.

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1 Introduction

It is a core interest of labour economics to understand better the relationship between labour supply and public policy. Over time, labour supply has continued to remain of great importance, particularly when it comes to women with children. This is because larger individual labour supplies can lead to higher incomes, independence from welfare programs and an increase in retirement income. As per the international statistics, female employment generally decreases following childbirth, and is particularly low for mothers with pre-school aged children. For instance, within the Organisation for Economic Co-operation and Development (referred to OECD hereafter), only 56 per cent of women were working in 2005, compared with nearly 75 per cent of men (OECD, 2007). More specifically, the employment rate of women with children aged between 6–16 years was 63.2 per cent, 58.2 per cent for those with children aged 3–5 years and only 51.1 per cent for the mothers of children younger than 2 years of age[1].

The Scandinavian countries are among those countries with highest maternal employment rates and some studies advocate that one factor driving this finding is the provision of childcare (e.g. Gupta et al., 2008; OECD, 2007). It is then an important question from a public policy viewpoint whether the increased provision of public childcare increases female employment and hence tax revenue through increased employment. In principle, it is possible that the increase in the provision of formal childcare does not actually induce women to shift from not working to working, but merely encourages working mothers to shift from informal to formal childcare arrangements, in a process referred to as crowding out. Another important question is whether adjustment occurs at the intensive margin, inducing women at the mean to work more hours or to shift from a part- to a full-time job. Both changes would also generate higher tax revenues for governments.

An increasing number of studies have estimated the causal effect of childcare provision on female employment. However, the overall evidence remains quite mixed, and questions remain about whether the expansion of childcare for pre-school-aged children actually leads to an increase in the female labour supply. This is particularly the case for the mothers of children aged 1–2 years. In the present study, we present new evidence on the causal effect of childcare on maternal labour supply using the increased provision of public childcare for the youngest children (those aged 1–2 years) through a nationwide reform in 2002 in Norway. This particular reform increased childcare coverage within a relatively short time from under provision to full

[1] The numbers are for the EU countries.

demand coverage, that is, 80 per cent of the birth cohort having a childcare place. The target group we focus on are mothers whose youngest child is aged 2 years or younger. We consider outcomes measuring labour supply at both the extensive and intensive margins.

Our analysis contributes to and extends the literature in three ways. First, to address the endogeneity issue when estimating the causal effect of childcare provision on the labour supply, we make use of a childcare reform that creates variation in the expansion of childcare across all municipalities in Norway. Because the actual change in childcare slots may be correlated with labour market conditions that also affect labour supply outcomes, we exploit the pre-reform level of childcare coverage within a municipality as a predetermined variable. We present a reduced-form estimate that exploits, as we show, that the pre-reform level of childcare coverage within a municipality captures the intensity of the response at the municipality level. Second, we present estimates of the responses in the labour supply at both the extensive and intensive margins by using information from the labour force survey on both the actual and contracted hours of work. This is of benefit because it is well known that employment rates of mothers in national and international statistics may be overstated, as women may have a working contract designating hours to be worked, but may not actually work for positive hours given job-protected parental leave rights. To reflect this, we employ information on the actual hours of work to measure precisely whether mothers shortly after childbirth are on contracts and actually working. During this period, mothers may also be on paid or job-protected parental leave, hence not actually working, or may actually work reduced hours that fall below their contracted hours. The labour force survey we use is sufficiently large and of very high quality to allow us to focus on the target group of mothers. Third, instead of focusing on mean post-reform effects as in most of the existing literature, we test for heterogeneous effects in both the short and long term, and across sub-groups of mothers who may vary in their responsiveness to reform (post-maternity leave, with more than one child).

Our main result is that mean maternal employment significantly increased for our target group of mothers. However, the evidence on the contracted hours of work is weakly significant. The adjustment mainly is in the short term (1–3 years) after the reform. The quantity of the employment effect corresponds to a 2–2.5 percentage point increase in employment with the effect in contracted hours amounting to slightly more than 1 hour per week when childcare coverage increases by 10 percentage points[2]. We find no significant effects on working long

[2] The calculated effects are based on an average increase of childcare coverage by 10 percentage points. The extensive margin effect is a raw measure of the increase in the employment rate. The pre-reform average employment rate for all mothers with at least one child below the age of 3 years is 72 per cent. In relative terms,

hours and actual hours of work. Nonetheless, we also identify substantial heterogeneity across sub-groups of mothers in the affected outcomes and the timing of the effects. When we include only mothers after the expiry of job-protected maternity leave—usually taken during the first year post-childbirth—in our estimation sample, the estimated reduced form effect on employment increases. In addition, the effect on actual hours substantially increases quantitatively and becomes significant. For mothers with more than one child, we also find significant short and long-term adjustments in employment and long-term effects on the actual and contracted hours of work.

The remainder of the paper is structured as follows. Section 2 reviews the literature on public childcare and maternal employment. Section 3 describes the institutional background of the designated childcare reform. Section 4 outlines the empirical strategy and Section 5 describes the data and presents some descriptive statistics. Section 6 presents the empirical results along with several robustness tests. Section 7 provides some concluding remarks.

2 Public childcare and maternal employment

According to a standard static labour–leisure choice model, we expect that the expansion of public affordable childcare leads to positive labour supply responses at the extensive margin because of the reduction of opportunity cost. This effect may be more pronounced with the expansion of childcare for children aged 1–2 years in comparison with older age groups. This is because care intensity is highest among these children. In addition, if we consider the dynamic labour supply effects through detachment from work post-childbirth, then expected employment responses may be even larger for mothers of younger than for those with older children. At the intensive margin (conditional on working), the predicted effect may be very heterogeneous. The effect may also be nonlinear in the hours of work.

Of the existing empirical studies concerning the effects of the provision of public childcare on the female labour supply, most focus on the effects of childcare provision for pre-school children aged 3–6 years, with few studies examining the effects of childcare provision for children aged 1–2 years. For its part, the overall evidence on the effect of childcare for 3–6-year-olds on the maternal labour supply is surprisingly mixed, and the results are often difficult to interpret. This is because earlier studies often depended on restrictive assumptions

the increase in employment rate is then 3 per cent. At the intensive margin, the 1 hour per week increase includes increases from zero hours. The corresponding marginal effect at the intensive margin conditional on positive hours is then approximately 0.7 hours per week.

to identify the direct effect of childcare provision (e.g. Ribar, 1992, Kimmel, 1998). This is important because if these restrictive assumptions do not hold, the estimated effect is contaminated by other factors that potentially affect both childcare provision and the female labour supply simultaneously.

It has been especially challenging to find convincing exogenous variation that affects childcare provision, but not the female labour supply. More recently, some studies have exploited the variation in the number of childcare places for children aged 3–6 years as a proxy for treatment intensity, across regions and time, as exogenous variation. This strategy is similar to the approach taken in Card's (1992) seminal work in which regional and time variation of the introduction of minimum wages was used to estimate the effect on youth wages. Specifically, these studies exploit a number of public childcare expansionary reforms targeting the 3–6-year-old age group among pre-school children in various time periods since the 1960s. These studies are informative not only about the female labour supply, mostly at the extensive margin, but also about now widely debated policies with the goal of universal childcare, with quite high quality and at affordable prices. Past studies cover Canada, the United States, Germany and Norway, with most showing a positive effect on maternal employment, including Baker, Gruber and Mulligan (2008) and Lefebvre and Merrigan (2008) for Canada, Bauernschuster, Hener, Rainer (2016) for Germany and Cascio (2009) for the United States. The exception is a Norwegian study exploiting a universal childcare reform in 1975 by Havnes and Mogstad (2011)[3], which revealed that the magnitude of the maternal employment effect varied.

It is interesting to note that the settings in past studies differ according to the specific characteristics before and after the respective reforms, which may explain the differences in the significance and size of the responses. Early reforms used in these studies took place during times of relatively low childcare coverage and low maternal employment, at least during the pre-reform periods (Havnes and Mogstad, 2011). By contrast, subsequent reforms were typically conducted in an environment where it was already more common for women to have children in public childcare and work. As a result, reforms initiated at lower levels (in both childcare coverage and maternal employment) tend to show relatively large employment effects, whereas reforms at relatively higher levels tend to reveal smaller employment effects. An important aspect is also whether the crowding out of informal care through formal public care is reducing the employment effect.

[3] Fitzpatrick (2010) provides additional evidence of the small maternal labour supply effect in the United States using universal pre-kindergarten programs in the 1990s.

Responses may also be relatively low in regimes where the work–family balance is low, such as in most countries during the 1980s or earlier. Employment effects also tend to be lower with lower rates of childcare coverage after the reform. Interestingly, reforms tend to be more effective in terms of the maternal labour supply if it is the youngest child of the mother that is affected by the reform (see also the detailed discussion in Bauernschuster and Schlotter, 2015). An additional aspect that may diminish any employment effects is that these reforms affect mothers from the third year after the birth of their youngest child. Hence, detachment effects from work or the depreciation of human capital may reduce the observed employment effects.

Another aspect varying across reforms is how many hours of childcare are provided each day. If these cover only a small proportion of a full-time job, then the estimated labour supply effects are expected to be smaller. From the parental leave literature, we know that longer parental leave decreases both wages and the probability of a return to work (Ruhm, 1998; Lalive and Zweimüller, 2009; Ejrnæs and Kunze, 2013). However, studies in this stream of the literature do not typically consider the costs of childcare for pre-school children or the provision of universal childcare. This is important because childcare costs are typically substantial during the phase immediately following childbirth. It is then of particular interest to examine the maternal labour supply effects for that group of mothers where the youngest child is aged 1–2 years, as these mothers are potentially employed and transitioning from work to job-protected and paid parental leave before returning to work.

The literature examining the effects of childcare provision for children aged 1–2 years is quite limited. Among past studies, only studies by Baker et al. (2008) and Lefebvre et al. (2008, 2012) in Canada have exploited regional variation in a nationwide reform targeting the youngest children (though much broader age groups than in our analysis)[4]. See Table A1 in the appendix for details of selected studies on maternal labour supply responses to childcare expansionary reforms. These Canadian studies exploit regional variation in the growth rate of expansion using a differences-in-differences framework and identify positive maternal employment effects in response to a reform in Quebec.

A French study by Goux and Maurin (2010) has also found positive maternal employment effects for single mothers in response to the availability of a formal classroom-based learning environment for 2- and 3-year-olds. For identification, they make use of the highly discontinuous relationships between a child's exact date of birth and pre-elementary school eligibility. They also use data for children born between 1995–1997, when most children

[4] Another study is by Simonsen (2010) that exploits regional variation in the childcare price in Denmark for identification.

entered childcare at age 3 years, and only about a third entered early at age 2 years. Goux and Maurin (2010) found no significant effect on mothers in two-parent households, with the suggested interpretation being that two-parent households have greater access to informal care because of wider networks; therefore any crowding out may reduce the estimated employment effects.

Bettendorf et al. (2015) considered the introduction of a childcare law affecting children aged less than 4 years in 2005 in the Netherlands, also using a differences-in-differences framework, and identified a significant and positive effect on maternal employment. Similarly, Nollenberger et al. (2015) exploited an expansionary reform of subsidized public childcare in Spain and estimated the employment effects for mothers of 3-year-olds. Compared with the Netherlands and France at the time of the study (1987 to 1997), female employment rates in Spain are extremely low: 29 per cent compared with 70 per cent and more. Using differences-in-differences estimates, Nollenberger et al. (2015) estimated a significant positive effect on maternal employment that appeared larger than that for comparable studies in the Netherlands. Lastly, Bauernschuster and Schlotter (2015) used the introduction of a legal claim to kindergarten in Germany during the 1990s to study the effects of public affordable childcare provision on maternal employment and hours of work. They find quite a large positive and significant effect on employment[5].

It is interesting to consider the effects at the intensive margin alongside the evidence on the effects at the extensive margin. This is particularly the case for countries where regulations regarding working hours allow workers to choose part-time work and where firms offer part-time workplaces. As the international statistics show, considerable variation in the incidence of part-time work exists across countries. Generally, women, particularly mothers with pre-school children, are by far the main group working part-time. As an example, 60 per cent of women work part-time in the Netherlands, a particularly high incidence of part-time work internationally. Bettendorf et al. (2015) drew on this to reveal the effect at the intensive margin for mothers of new-borns to 3-year-olds through the reduction of the parental fee for childcare of 1.4 hours per week. Bauernschuster and Schlotter (2015) also identified the positive and quite substantial effect of 2.5 hours of work per week. In summary, the previous evidence reveals positive effects for maternal employment at both the extensive and intensive margins for mothers of children aged younger than 3 years[6].

[5] Noteworthy, in this study, childcare is only provided within short and restricted hours, normally between 08:00 and 12:00.

[6] Herbst (2017) also finds positive employment effects for childcare reform during WWII for the United States.

Given the vast interest in public debate on the introduction and expansion of tax-financed universal childcare for younger children, there is a great demand for more research on countries introducing universal and affordable high-quality childcare to derive evidence-based policy advice. Outcomes related to actual hours of work in addition to contracted hours and possible transitions of women working short part-time to longer hours are particularly under researched in this literature. In our analyses, we emphasize female labour market responses to the reform at both the extensive and intensive margins. We use information on the actual hours of work to allow us to measure exactly whether mothers shortly after childbirth have a work contract but are on (paid or job-protected) parental leave and hence not actually working. Mothers may also use this right to reduce the hours of work and are therefore actually working fewer than their contracted hours. Additionally, we emphasize the short- and long-term effects of childcare reform, as well as the heterogeneity across sub-groups of mothers.

3 Institutional background

In Norway, female employment rates were relatively high during the 2000s, which is our period of observation. As shown in Figure 1, the female employment rate was only 55 per cent in 1975, and increased rapidly to 70 per cent by the end of the 1980s. Since 1998, female employment rates have moved even higher, to between 72–75 per cent. Female employment rates in Norway, similar to other Scandinavian countries, are among the highest in the world. The figure illustrates this pattern where we compare Norway to Germany, the United Kingdom and the United States.

Norwegian policy has long placed a strong emphasis on family–work balance through a menu of family-friendly policies. Starting in the 1970s, the main policies to achieve these goals have been the introduction and expansion of public childcare and parental leave complemented by a cash-for-care policy. In this section, we describe the institutional background of the 2002 childcare reform we exploit in our analysis of the maternal labour supply effects. We also present further background on other child- and labour market-related policies that have not changed during the period we study, which is a key assumption to our estimation strategy.

Figure 1 here

3.1 The 2002 childcare reform

The childcare reform in 2002[7] that we consider followed the program of expansionary childcare reform starting in 1975 that achieved its goal of a high supply of childcare for preschool children aged 3–6 years. The Kindergarten Act in June 1975 provided the first large-scale public funding programme in Norway to expand childcare and make universal childcare available. The childcare coverage rate for 3–6-year-olds increased from 10 per cent in 1975 to 40 per cent in 1985 and approximately 60 per cent in 1997. During this period, childcare provision for children aged 1–2 years was much less common, particularly before 1987[8]. In 1997, the coverage rate for these children was approximately 40 per cent, and there were many parents who applied for but did not receive places in public childcare[9]. Hence, there was a surplus demand for public childcare.

We exploit the 2002 reform that was part of a large programme creating approximately 48,000 childcare places for 1–2-year-olds within a 6-year period. In Norway, a birth cohort is about 55,000 children. The main goal was to guarantee the availability of affordable childcare to all parents. Full coverage was defined by the government as a childcare coverage rate of 80 per cent, that is, 80 per cent of a birth cohort will have a place in a kindergarten within each municipality. To create a measure of childcare coverage in public subsidized kindergartens for each municipality and year, we use the municipality registers that report accurately the number of children with a childcare place and the number of children born. The number of children in childcare is the number of children actually enrolled at a public kindergarten either full or part time. Since 1998, the municipality register includes all kindergartens owned by the municipality, the county or the central government. Children with a childcare place in private kindergartens that receive subsidies are also included in the counts. The age of the child is measured at the end of the year. The childcare coverage rate is then the ratio of the number of children in public subsidized kindergartens for each age cohort, year and municipality, and the number of childbirths in the corresponding cell.

Figure 2 depicts the average childcare coverage rate for 1–2-year-olds when we take the mean across all municipalities for each year. For comparison purposes, we include the corresponding coverage rates for children aged 3–6 years. As shown, the childcare coverage rate was quite stable, at around 40 per cent, before the reform. Since the reform in 2002, there

[7] Part of this section is based on the governmental report - Kunnskapsdepartement (2012) NOU 2012.1
[8] OECD (1999).
[9] This has been documented in several reports, e.g. Kunnskapsdepartementet (2008).

was a gradual increase until it reached the goal of full coverage of 80 per cent in 2008; the coverage rate then again stabilized at this new and higher level. In our empirical analysis of maternal labour supply responses, we exploit this substantial quantitative expansion of childcare.

Figure 2 here

The reform background is that in 2002, the SV (*Sosialistik Venstreparti* or Socialist Left) Party put forward an agreement together with other large political parties to propose a second childcare reform following the first in 1975. The goals of the reform were threefold according to the reform proposal from 11 June 2002. First, all municipalities would have the responsibility of providing childcare to every pre-school child aged older than 1 year. Municipalities would also need to organize the application process for kindergarten places, that is, the handling procedure of applications by parents for a place for their child. This is an important aspect, given that before 2002, not all parents who applied for a childcare place in public childcare received an offer, and hence, queues existed for kindergarten places. Second, the law introduced an annual deadline of the 1st September and (formal) waiting queues. Lastly, at the same time, childcare fees were to be decreased, partly through a reduction of the full price for siblings.

The authorities implemented these changes quite rapidly. In spring 2003, the Norwegian parliament agreed on lower prices/fees for parents for a childcare place and the expansion of childcare coverage through public funding. In May 2004, a maximum price was introduced for a kindergarten place, and a reduced price for a kindergarten place for siblings. This resulted in a reduction of the price for a childcare place by 34 per cent in comparison with the average price in public kindergartens in 2002. The maximum price was less than 260 Euros per month (2,330 NOK) in 2008[10].

The Childcare Act (*Barnehageloven*) from 17 June 2005 contained the changes initiated in 2002. In addition, a focus was set on the increase and guarantee of high quality of childcare. This included agreement on a number of quality criteria, such as number of childminders per child. In our study, we have no information on childcare institution characteristics and hence, we do not consider the quality of childcare. Since 25 April 2008, this law guarantees the right to a childcare place within the municipality where the child lives. As shown in Figure 2, while

[10] See the Childcare Act (*Barnehageloven*) nr. 52, p. 7.

coverage for 3–6-year-olds was already high (but less than 80 per cent), new places were also built for this older group in this process in some municipalities.

Municipalities attained the goal of full childcare coverage through cooperation with private parties and entrepreneurs. Within 5 years after 2003, they built an additional 48,200 full-time kindergarten places. The state budget allocated to kindergartens also increased by 177 per cent, from 7.8 billion Norwegian Krones in 2003 to 21.6 billion in 2008. The expansion and construction of new kindergartens was financed through three main sources: the state budget, the municipality budget and parental fees for kindergarten usage. Before 2003, only the state budget finance was regulated by law. At the municipality level, a new regulation was added such that kindergartens had to be equally treated economically relative to the local budget. The law also established equal subsidies to private and public childcare institutions that satisfied the federal provisions; previously, private institutions had been awarded only 85 per cent of the subsidy rate offered to public institutions. All acknowledged kindergartens in a municipality have the right to public subsidies of a regulated amount.

A typical childcare place in Norway is full time, Monday to Friday, with opening hours from 07:30 to 16:30. This corresponds to normal full-time working hours. A kindergarten has groups for children aged 1–6 years. Since 1997, the school starting age in Norway has been 6 years of age (the calendar year when the child turns 6). Groups with 1–2-year-old children must have more childcare providers and pedagogues per child than those with children aged 3–6 years. Kindergartens are open year round except July, which is the usual summer holiday month in Norway. New intakes usually take place once a year on 1 September[11].

This reform overview highlights several features that make it particularly suitable for studying the effects of an increase in the number of childcare slots for 1–2-year-olds on the maternal labour supply at both the extensive and intensive margins. First, it is a large-scale reform providing universal public childcare that affects mothers soon after childbirth. Second, the childcare is full time, enabling mothers to combine family and work, including jobs with longer working hours (longer part-time or full time). Therefore, we may expect labour supply effects at both the extensive margin (new labour market entries) and the intensive margin (transitions from shorter to longer working hours or full-time work). In our analysis, we focus on the quantity of childcare slots. Our empirical analysis thus estimates the total effect of the reform, and we cannot account for changes in quality. The price effect is part of the total effect and partly captured by time effects.

[11] In the empirical analysis, we use yearly not monthly data on childcare slots and the age of the child.

3.2 Possible employment effects in the childcare sector through expansionary reform

The kindergarten sector is dominated by female workers. The expansion of this sector, hence the demand, could generate positive incremental employment effects at the aggregate level by shifting work traditionally conducted unpaid by mothers at home or in the informal sector to the paid public sector. However, we expect them to be small for our target group of mothers. In Table 1, we present for selected years the number of employees in kindergartens, the percentage increase over 5-year intervals and the percentage of employed women in the kindergarten sector relative to total female employment in the Norwegian economy. Two trends are noteworthy. As expected, following the childcare reform in 2002, there was a substantial increase in employment in the kindergarten sector. From 2000 to 2005, employment increased by 22.8 per cent and, by 2010, it increased by an additional 35 per cent. For comparison, the 1975 Kindergarten Act created 14,600 new workplaces, while the 2002 reform a bit less than 24,000 new workplaces.

This gives us an idea of the relative size of the reform, but also that care for 1–2-year-old children is more labour intensive. Aggregate employment numbers for women also increased in the economy overall during this period. The last column of Table 1 details the proportion of women working in the kindergarten sector after subtracting men working in the kindergarten sector[12]. As shown, female employment in this sector increased by 0.5 percentage points from 2000 to 2005 and by 1.2 percentage points from 2005 to 2010. The labour supply effects we are identifying by use of the reform will contain this added female worker effect, if there are any mothers with young children among these.

Table 1 here

4 Empirical strategy

It is challenging to estimate the effect of childcare on the female labour supply. The main reason is the simultaneity problem in the maternal employment regression. Does childcare expand first

[12] During the same period, the employment of men in the kindergarten sector also increased quite substantially, which may be the result of the positive anti-discrimination policy in the recruitment policy introduced in 1998, titled The Positive Discrimination Act for Men (*Forskrift om saerbehandling av menn*) in Kindergarten and Other Educational Sectors Where Women Form a Majority.

and then lead to an increased maternal labour supply, or does maternal labour supply increase first for some other reason and then as a response, childcare provision is improved?

In order to estimate the direct effect of childcare on maternal labour supply, we exploit the expansionary reform in 2002 of public childcare for the group of 1–2-year-old children. Our main estimation approach is a reduced-form estimation approach where we exploit the regional variation in childcare provision in 2001, the year before the expansionary reform, as a predetermined variable[13]. In order to provide a convincing empirical base for our identification, we show that the pre-reform regional variation in childcare provision is positively correlated with the differential increase in childcare expansion (see Figure 4.1). As the outcome variable, we use different measures of labour supply at the intensive and extensive margin, including the number of contracted and actual hours of work. We present estimates of the short- and long-term effects of the reform where we divide the post-reform period into the short-term post-reform period 2002–2004 and the long-term post-reform period 2005–2009.

The main regression we estimate is specified as follows:

$$y_{i,t} = \beta_0 + \beta_1(Postshort_t) + \beta_2(Postlong_t) + \beta_3\left(Precoverage_{i_{2001}} * Postshort_t\right) + \beta_4(Precoverage_{i2001} * Postlong_t) + \beta_5 * X_{it} + \beta_6 * M_i + \epsilon_{i,t}, \quad (1)$$

where *i* indexes individuals and *t* time periods, and *y* represents the labour supply outcome variable. We present estimates at the extensive margin on employment and working for long hours (i.e. longer than 18.5 hours), and we also provide intensive margin results on both actual and contracted working hours per week. The variable *Precoverage* measures the pre-reform coverage rate in 2001 in the municipality where the mother is living. This variable is a predetermined factor capturing the intensity of the municipal response to the reform. The indicator variable $Postshort_t$ takes a value of one if the observation is from 2002 to 2004, zero otherwise, and the variable $Postlong_t$ takes a value of one if the observation is from 2005 to 2009, zero otherwise. X_{it} includes individual characteristics that we add as additional control variables for the main determinants of the female labour supply. These are age, age squared, education, marital status and the number of children under the age of 16 years. M_i includes municipality fixed effects for the more than 400 municipalities in our data. Inclusion of

[13] Our empirical strategy draws on Løken, Lundberg and Riise (2016) to identify the causal policy effects.

municipality fixed effects ensures that we control for municipality characteristics that may be correlated both with the level of childcare provision before the reform and the outcome variable. $\epsilon_{i,t}$ captures idiosyncratic shocks to labor supply. We present robust standard errors clustered at the municipality level. This is to account for shocks that are common to all mothers to be correlated between mothers living in the same municipality.

Depending on the nature of the labour market outcome variables, we use several different regression models. We first estimate a linear probability model with the binary outcome variables of employment status and working long hours. We also estimate Tobit models for actual and contracted hours, as these variables contain zero values[14]. We test our specification for whether the reform generated heterogeneous effects over time depending on the phase-in schedule of the policy adjustment. It is an empirical question whether the short- or long-term effect is stronger. A hypothesis is that the strongest labour supply responses arise after 2005 when the reform fully unfolds and childcare coverage rates are above 50 per cent.

The main coefficients of interest are β_3 and β_4 in eq. (1). These are the coefficients of the continuous variable measuring the pre-reform childcare coverage rate of the municipality the mother is living in and the indicator variable for the post-reform period. Note that the pre-coverage variable is defined between zero and one. We interpret the coefficients as the intention-to-treat (ITT) effects, or the reduced-form parameters, of the reform on the outcome y in the short term, 2002–2004, and the long term, 2005–2008. This interpretation is similar to previous work, such as Baker et al. (2008) on the overall average effect. We choose a reduced-form specification that is linear in the pre-reform childcare coverage rate, *Precoverage*.

Our approach also controls for individual characteristics that capture important control variables in labour supply regressions. We control for time effects for common time-varying changes. In addition, we control for municipality fixed effects. As these capture demand factors likely correlated with the pre-reform childcare coverage rate in a municipality in 2001 as well as maternal employment outcomes, our results are not biased because of these factors.

Figure 3 here

To illustrate the variation in the *Precoverage* variable that we exploit, Figure 3 provides the geographical map of Norway with municipalities and mean childcare coverage rates across municipalities. The first panel shows the coverage rates in the pre-reform period in 2001. The

[14] For example, 27 percent of the full sample of mothers report zero contracted hours for the period analysed.

municipalities in light grey have coverage rates between 10–28 per cent. The municipalities with larger than 53 per cent coverage are in dark grey. The mean coverage rate during this period is 40 per cent. We see from this map that low and high coverage rates were quite scattered across areas and variation is considerable.

Figure 4.1 here

All municipalities could in principle apply for public funding and expand childcare. In the data, there is evidence that municipalities with low pre-reform coverage levels are those growing more strongly than those with relatively large pre-reform childcare coverage levels. The correlation is graphically presented in Figure 4.1, which clearly reveals a negative relationship. We can also see in the data that the expansion of childcare places for the youngest group of children, the 1–2-year-olds, was an independent expansion to the childcare places for 3–6-year-olds, and there is no significant correlation between them, as shown in Figure 4.2.

Figure 4.2 here

One possible concern is that we may be confounding the estimated childcare effects with other reforms or changes during the period we consider. Of these, two policies were in place in Norway that could potentially affect the labour supply and the take up of childcare, both introduced or reformed in the 1990s: parental leave and cash for care provision. For our particular period of analysis, we identify no other significant reforms or breaks in trends that could be of concern for our estimations.

The duration of paid parental leave is likely to affect the return to work decision and therefore the demand for childcare. However, during our period of analysis, all women were eligible for approximately 1 year of paid leave and there were no additional reforms. Since 1993, all working parents in Norway have been entitled to 52 weeks leave with 80 per cent wage compensation (alternatively, 42 weeks with full compensation). To increase the involvement of fathers, an amendment in 1993 reserved 4 weeks of the leave for fathers. Since 2005, the father's quota, i.e. leave duration, has gradually been extended to 10 weeks. The latest figures show that on average, fathers take 10 weeks of paid leave, typically at the end of leave, and mothers, 46 weeks (Samfunnspeilet 2012/1). Following parental leave, the most common pattern is that both parents return to work. Fathers remain in full-time work typically and do not adjust labour supply. In Norway, full-time work is 37.5 hours.

Another family-friendly policy was introduced in 1998, being the "cash for care" program. Parents of children who do not make use of a full-time childcare place that is publicly subsidized receive approximately 260 Euros per month for children aged 1–2 years[15]. Starting from 1 January 1999, cash-for-care benefits were extended to children aged 2–3 years, which was later reduced[16]. Hence, there were no other changes during our analysis period for this policy that could confound our estimates.

5 Data

Our data are extracted from the Norwegian Labour Force Survey (LFS)[17], a quarterly dataset containing a representative sample of the Norwegian population for every year. We use the information for the period 1998–2009. The LFS provides information on individuals between the ages of 15–74 years regarding their labour market activities, demographic characteristics and education. The survey collects information from a representative sample of residents randomly selected from all municipalities on the basis of a register of family units. The sample consists of about 12,000 family units (24,000 persons) each quarter. Each family member is in the survey, answering questions about their situation during a specified reference week. Demographic data are compiled from the Central Population Register, and data on education are supplemented from the register of individual data collected by Statistics Norway from the educational institutions. The Norwegian LFS provides a repeated cross-sectional data source.

For each year, we use individual information on demographic variables such as gender, age, number of children aged 16 years and younger, age of the youngest child, marital status, education and the municipality of residence. Information in the Norwegian Labour Force Survey is of high quality and is the main source of information for the national accounts statistics. The main purpose is to provide information on individual employment and unemployment, and on the labour force participation of different groups of the population; these are the main indicators reported regularly by the government. A key advantage for our analysis is that the database allows us to analyse labour supply at both the extensive and intensive margins using both actual and contracted hours of work.

Sample selection

[15] Cash for Care Act of 26 June 1998 Nr. 41 for Parents of Infants.
[16] On 12 Dec. 2012, the Norwegian parliament agreed that cash for care would only be paid for children aged 1–2 years.
[17] The LFS is available from the Statistical Office of Norway (*Statistisk Sentral Bureau* or SSB).

Our main analysis sample contains mothers between the ages of 17–49 years whose youngest child is aged 2 years or less[18]. This is the largest group potentially using childcare for the youngest children and immediately responding to the expansionary reform in 2002[19]. As our observation window for the main analysis is 2001–2009, we exclude some mothers potentially benefiting later from the reform, but whose children are older[20]. Our final estimation sample contains 27,577 mothers for the period 2001–2009. If we restrict our sample to the year 2001, we have 3,439 mothers. We use 1 year before the childcare reform and 7 years after the reform to estimate the short- and long-term effects given the reform in 2002. In a robustness test, we also make use of the period between 1998–2001 to test using a placebo scenario whether the treatment and control areas have a common trend leading to the reform year in 2002.

In the empirical regression analyses, we start with estimations of our model as depicted in equation (1) on the largest group of mothers with their youngest child aged 2 years or younger; this is the immediate target group of the reform. Additionally, we test whether the effects are heterogeneous across sub-groups. The first sub-sample excludes mothers who report that they are on maternity leave and work zero hours. Most mothers in Norway take 1 year of job-protected and paid parental leave[21], and are eligible for an additional year of unpaid leave. By excluding all on maternity leave, we can test whether the remaining group that is no longer on job-protected maternity leave is more responsive in terms of labour supply adjustment. Furthermore, we analyse the sample of mothers with more than one child to see if the reform differentially affects mothers facing different time constraints and parity. In addition, a focus on mothers with mostly two children may capture that group of mothers who will likely not have more children and therefore make an upward adjustment in the labour supply more probable.[22] Additionally, we select a sub-sample of mothers with more than one child who are not on maternity leave, equivalent to a period of working zero hours but having positive contracted hours.[23].

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[18] We assume that fertility is completed before the age of 50 years for most women. This is confirmed in the data, where no women older than age 50 years reported having a child younger than the age of 2 years. We also used a different age group 25-49 to see if our results are sensitive to age group selections. We do not find that the results change much due to different age groups. There are very few mothers who were below age of 25 in our sample.

[19] We exclude the self-employed.

[20] For example, these are mothers whose youngest child is aged 3 years in 2006. Note that we cannot follow mothers over time in the cross-sectional LFS.

[21] We acknowledge that this selection is to some extent endogenous. Alternatively, we selected those with the youngest child aged 1–2 years. The results are robust and available on request.

[22] In our sample, the majority of mothers who have more than one child have two children for at least the period we observe.

[23] We also tested whether excluding big cities affects the results, and they remain unchanged.

Outcome variables

Our main measure of labour supply is employment status—whether a person is employed—reported in the labour force survey based on contracted hours. Generally, in the Norwegian context, full-time work is 37.5 hours per week[24]. The LFS contains continuous variables for actual and contracted hours of work. Using actual hours is particularly important to measure labour supply of mothers with young children. Note that as long as a person is on job-protected (or paid) parental leave, the person is formally employed, and actual hours need to be measured in order to assess employment at the intensive margin. This is also because in Norway, part-time work is quite common and many women reduce actual hours of work post-child birth, but not necessarily contracted hours, given that within a certain period, people on parental leave have the right to reduce (actual) hours of work.

To measure labour supply more accurately at the intensive margin, we use both actual and contracted hours of work. We specify actual self-reported working hours as one outcome variable. This variable measures, at the individual level, total actual working hours during a typical working week, i.e. including overtime and excluding absence from work[25]. Based on actual hours, we generate our second outcome variable, an indicator variable for long hours, which we define as 18.5 hours per week (more than half the typical full-time work hours). This variable captures employment status at a higher level of intensity. In addition, we make use of contracted working hours. This variable captures the weekly number of working hours in the working contract. These four variables are the outcome variables that we analyse, and they provide us with a very detailed picture of labour market activity at both the extensive and intensive margins for our target group of mothers with young children

Individual control variables

We use the age of the mother, which is the age at the time they completed the survey, and years of completed education. In the survey, the educational level is reported in categories, which we

[24] In the labour force survey, employed persons are persons aged 15–74 years who performed work for pay or profit for at least 1 hour in the reference week, or who were temporarily absent from work because of illness, holidays, parental leave, etc.

[25] Note, persons absent from work are generally not included in the calculations of actual working hours per week. We top code actual weekly hours at 60 hours/week.

then transformed into number of years of formal education[26]. We define marital status to distinguish those married from those single or divorced/widowed. We also use information on the number of children in the household. We count all children aged 16 years or younger to capture dependent children in the family. Following the literature on maternal labour supply, these individual characteristics are the commonly used variables determining female labour supply. To merge our data with municipality level information on childcare coverage, we also extract information on the mother's municipality of residence.

Childcare coverage

To construct the childcare coverage rate within a municipality in a given year, we extract information on the number of children with a childcare place from 1998 to 2009 and childbirths from 1995 to 2009 from the municipality registers. As a result, we can distinguish 448 municipalities. We then calculate the childcare coverage rate as the ratio of the number of childcare slots in a given year for children at a given age, say 1–2 years, to the number of children in the respective birth cohorts.

Summary statistics

Table 2 here

Table 2 provides summary statistics for the main sample and the three sub-samples measured in the pre-reform period 2001. Overall, we observe that the employment level is high. Column 1 reports the means for the main sample showing that the employment rate is 72 per cent, and that 55 per cent work longer than 18.5 contracted hours. Contracted hours at the mean are close to 22 hours per week. We see from the comparison of contracted and actual hours that on average, mothers with children aged 2 years and younger actually work fewer than their contracted hours (only 11 hours, including zeroes). The discrepancy is quite large and mainly reflects interruptions due to maternity leave when the worker is employed but not working, or

[26] The category and years of schooling matches are as follows: "7 years" = "elementary 7 years of schooling"; "10 years" = "middle school 10 years of schooling"; "12 years" = "12 years of schooling, high school"; "13 years" = "13 years of schooling, higher level of high school"; "14 years"="technical college"; "16 years"="university and college"; "18 years"="higher university and college"; and "20 years"="researcher education".

reduced hours of work during leave. It could in principle also be due to health reasons; however, this is less likely.

We confirm this when we calculate the corresponding averages for the sub-sample of mothers, excluding those who report maternity leave in column 2; this group amounts to 74.1 per cent in our sample. Now the difference between contracted and actual hours decreases to 3 hours (18 contracted versus 15 hours of actual hours of work), compared with a difference of 11 hours in the main sample. It is likely that women on maternity leave do not reduce contracted hours compared with before childbirth, and since they are not working during maternity leave or reducing hours of work, this discrepancy arises. We now also see a lower probability of being employed, at 63 per cent. This means that there is a group of women not working following the expiry of job-protected parental leave.

In the empirical regression analysis, we run separate regressions on the main sample of mothers and the sub-group of mothers, excluding those on maternity leave. This allows for the fact that mothers still on maternity leave (paid or unpaid) may respond differently to the childcare reform. When we look at the summary statistics for mothers with more than one child, the means are quite similar, although employment, the incidence of long hours and both actual and contracted hours are slightly lower. This could reflect increasing time constraints before the childcare reform.

To look more closely at the distribution of contracted hours of work, we present categorical means in hour brackets. For example, for our main sample, the average weekly working hours for mothers who work low (less than 18.5) hours is 12.37 per week. Those who work more than 37.5 hours per week on average work 39.57 hours, reflecting average hours for full-time workers. There is not much difference between the main sample and the three sub-samples in terms of the categorical means. The sub-samples where we exclude mothers on maternity leave tend to work slightly more than the groups including them. We observe quite large differences in the extensive margin between the sub-samples. For example, mothers without maternity leave options are less likely to work by 10 percentage points compared with those who have the maternity leave option; this difference remains for mothers with more children. It is noteworthy that mothers work quite long part-time hours, typically more than 18.5 hours overall. For those who work positive hours, around 85 per cent of mothers of young children work more than 18.5 hours, and 48.5 per cent work full time[27]. Beneath the summary statistics for the outcome variables, we report the means for the individual control variables. As

[27] The shares of each hour category are not provided in Table 2.

shown, the target group of women are on average 31 years of age, which is relatively young compared with the employed population. On average, they also have 14 years of education and 1.88 children, and hence, most have two children.

6 Empirical results

In Table 3, we present the results of the reduced-form estimates of the effect of the childcare reform, or the intention-to-treat (ITT) effects, for our main sample of mothers. The short-term ITT is the effect of the period 2002–2004 and the long-term ITT is the effect of the period 2005–2009. In separate columns, we report the results for the four outcome measures of labour supply. In all of the regressions, we control for individual characteristics, time effects and municipality-specific fixed effects. The complete regression results are presented in appendix Table A2. These reveal that the individual characteristics have the expected signs. Most coefficients are significant and capture part of the variation in the outcome.

For the main sample, we find a strongly significant effect of the reform on employment in the short term and a weak significant effect in the long term, and only a weak significant effect on contracted hours of work. The ITT effects for the outcome variables working long hours and actual hours are not significant. Hence, the strongest employment effects occur 1–3 years after the reform. This we may refer to as a phase-in effect when the reform started to unfold, and the percentage of children aged 1–2 years enrolled in childcare was starting to increase from a mean of 40 per cent, as shown in Figure 2. The point estimate of the coefficient is not only more precisely estimated, but also somewhat larger in the short than in the long term. In the short term, a 10 percentage point increase in childcare coverage increases mothers' employment by 2.47 percentage points from a base of 72 per cent, which is the pre-reform mean employment level in this group. This translates into a 3.4 per cent increase.

In the long term, the corresponding calculated effect is 2 percentage points. In addition, we find positive effects on contracted hours, but not on actual hours, are only significant at the 1 per cent level. The coefficient is 10.434 in the short term and 7.223 in the long term. This translates into a marginal increase of 1.04 hours per week, conditional on all hours, including zeroes, and 0.68 hours conditional on only positive hours, which amounts to a 2.2–4.7 per cent increase in the short term. Comparing the coefficients in the regressions for contracted and actual hours, we note that the point estimate of the short-term effect is more than double the size in the regression for contracted hours. Hence, the point estimate of the effect on contracted hours is larger but not very precisely estimated.

Overall, this yields strong evidence in favour of an expansionary effect on employment, which suggests that at least some women were induced to enter the labour market. The evidence is strong because we include all women that were potentially directly affected by the reform. However, we include one group of mothers who are still on job-protected maternity leave and receive 80–100 per cent wage replacement through the government-mandated parental leave system[28]. Therefore, they are less likely to adjust labour supply at the intensive margin during this period. Hence, we may underestimate the effect on actual hours. The summary statistics in columns 1 and 2 in Table 1, together with the first regression results, support this hypothesis. Another hypothesis we explore is whether the responses are stronger for mothers with more than one child, for whom we have seen from the summary statistics that before the reform, they worked slightly fewer hours in terms of both actual and contracted hours.

Table 4

Table 4 presents the results when we re-run the regressions on the three sub-groups of mothers[29]. In the first panel, we report the results when we exclude mothers on maternity leave. We find significant effects on employment, strongly significant in the short term and less precisely estimated in the long term. Again, we find an effect on contracted hours in the short term that is only significant at the 1 per cent level. The new finding is that now the effect on actual hours is also significant, although only at the 1 per cent level. Comparing the size of the point estimates to the previous results in Table 3, it is noteworthy that all of the coefficients are larger, and that the coefficients in the regression for actual hours increase from 3.99 to 12.81 in the short term.

This is consistent with incentives induced by the institution of maternity leave. Mothers can take job-protected and paid parental leave. During this period, we do not observe any intensive margin adjustments in actual working hours; it is after this period we expect, if any, negative adjustments. This is confirmed by the sub-group estimates, where a 10 percentage point increase in childcare coverage increases actual hours of work by about 0.69 hours per

[28] The LFS does not contain information on wages.
[29] The complete regression results are in Tables A2 to A5 in the appendix.

week, or 2.4 per cent (= [0.69 ÷ 28.99] × 100). Alternatively, if we consider zero hours, the effect would be 1.28 hours per week, or 8.4 per cent (= [1.28 ÷ 15.23] × 100)[30].

We further explore whether the results change when we select on mothers with more than one child. A majority of these mothers may have completed fertility and may therefore respond more to the reform, thereby increasing the estimated effects. In the second panel of Table 4, we report the corresponding regression results for the four outcomes. All of the coefficients remain large when we look at the regressions for employment, actual and contracted hours, but the significance of the effects shifts from the short-term period, 2002–2004, to the long term period, 2005–2009. The employment effects show that a 10 percentage point increase in childcare coverage increased employment by 2.69 percentage points in the short term, and by 3.57 percentage points in the long-term; these are now highly significant. We also see strong effects in the long term on the actual hours of work. On the intensive margin and conditional on positive hours, the marginal effect of increasing the coverage rate by 10 percentage points is 0.66 actual hours per week, which implies a 2.5 per cent increase in weekly hours[31]. We also document an increase of 0.89 hours on weekly contracted hours, which is about a 3.1 per cent increase in the long term. Overall, this suggests that after the completion of fertility (that is, two children for most mothers), mothers readjust towards stronger labour force attachment at both the extensive and intensive margins. Interestingly, these effects take place long after the initiation of the reform.

Panel 3 reports the estimates when we also exclude mothers on maternity leave. The results largely confirm those found earlier, but the long-term effects are now strongly significant and even larger in magnitude based on the point estimates. The employment effects show that a 10 percentage point increase in childcare coverage increases employment by 3.2 percentage points, which is marginally significant, in the short term, and 4.6 percentage points, which is highly significant, in the long term. We also see stronger effects in the long term on actual weekly hours of work conditional on working, by almost 1 hour, or 3.6 per cent, for a 10 percentage point increase, as well as a significant effect on contracted hours by 1.1 hours, or 3.9 per cent, for a 10 percentage point increase.

6.1 Robustness check

[30] This is a marginal effect assuming the effect is evaluated for positive hours only. The average positive weekly actual hours of work before the reform for this group was about 18 hours.
[31] This is the marginal effect on working hours conditional on positive hours.

We present two sets of robustness tests supporting our results. First, we show that the reform does not affect fertility or the number of new-borns. This finding supports our interpretation that we identify the direct impact on the labour supply given the reform. Otherwise, we could hypothesize that some women became mothers because of the increased availability of childcare and the decreased opportunity cost of having children. Second, we test the core assumption underlying the consistency of our estimation strategy. Our interpretation of the results assumes that in the absence of the reform, mothers in low and high childcare coverage areas would have had parallel trends in the labour market outcomes. This is similar to the assumption of a common trend assumption for the treatment and control groups in a differences-in-differences estimation approach. To test this assumption, we run placebo estimations where we move the reform backwards.

Table 5 here

Table 5 reports the regression output when we use new births as an outcome for individual fertility. To measure the incidence of a new-born child in a given year for an individual, we use the age of the youngest child reported in the labour force survey[32]. This results in a dummy variable for each mother equal to one if there is a new baby born to the family in a particular year, and zero otherwise. We estimate the regression model as specified in eq. (1) and replace the outcome variable by the new-born indicator variable. We find that none of the coefficients, short or long term, are significantly different from zero. Hence, we reject the hypothesis that the reform had a significant effect on fertility.[33] This is reassuring for our results and their interpretation.

Table 6 presents the placebo test results on the main estimation sample and the three sub-samples of mothers. We created a dataset that is extended backwards to 1998 and restricted the period to 1998–2001.[34] We then moved the date of the reform artificially back by 2 years, i.e. from 2002 to 2000. This yields an artificial pre- and post-reform period. We can then test whether there are any systematic differences in economic trends related to female labour

[32] Note, there is no information on the occurrence of twins in the survey, so our new-born indicator does not reflect the situation where multiple children could be born in the same year to the same mother. We believe such instances are rare and do not reflect the rational fertility decisions of a mother, as twins are hardly planned.
[33] The mean of the dependent variable is 0.023 in the pre-reform period which confirms that also the economic effect is close to zero.
[34] This dataset covers only four periods and hence only a short pre-reform and post-reform period to estimate the short- and long-term effects. Extending the dataset back even further would run the risk of contaminating the estimated effects with other parental leave reforms before 1998. Clearly, this would violate the placebo results.

markets in our treated communities (the municipalities with relatively low childcare coverage rates in 2001) and control communities (the municipalities with relatively higher childcare coverage rates in 2001). We generate a new variable that measures the childcare coverage rate within a municipality in 1998. We then define a new dummy variable $Postshort_t$ equal to one if the year is 2000 and a new dummy variable $Postlong_t$ equal to one if the year is 2001. The dummy variables are equal to zero if the period is 1998 or 1999. Table 6 reports the coefficients of the interacted variables that give the ITT estimates. The point estimates of all of the coefficients are relatively small and not significantly different from zero. Hence, we find no significant treatment effects across all specifications and sub-samples; we also find that the estimated effects are all quite small in magnitude. This is strong evidence supporting the parallel trend assumption. These robustness test results are reassuring for the causal interpretation of our main results.

7 Discussion

The results of our study provide strong evidence in favour of the positive maternal labour supply effects of universal childcare reform expanding childcare for 1–2-year-olds. We predominately find short-term effects of the reform for the immediate target group of mothers, that is, mothers whose youngest child is aged 2 years or younger. We also find long-term effects of the reform for mothers with at least two children. In more detail, for the entire target group, we find significant effects on maternal employment and weak evidence on upward adjustments of contracted hours in the short term for the large universal childcare reform. Surprisingly, we find no adjustments in actual hours, or transitions out of short periods of part-time work (i.e. less than 18.5 hours of work per week). As we show, these effects are relatively small and insignificant because some mothers are on paid maternity leave and job-protected leave. This is confirmed when we condition our estimates on not being on maternity leave, as the size of the effects increases on employment and contracted hours, but most remarkably, on actual hours of work.[35]

Our results showing significant and positive labour supply adjustments in response to a large childcare reform financed out of tax revenue are consistent with comparable studies that

[35] This suggests that estimates based on contracted working hours give downward biased estimates because of measurement error. This measurement may be generally large for mothers with very young children and in institutional settings with rights to paid und unpaid parental leave. Measurement error problems are at least reduced by use of actual hours of work.

have analysed childcare reforms targeting the youngest children and documented positive effects on employment (Baker et al., 2008) and on employment and contracted hours (Goux and Maurin, 2010; Bettendorf et al., 2015; Nollenberger et al., 2015). Accordingly, our results contrast with previous findings by Havnes and Mogstad (2011) showing that the childcare reform in Norway in 1975 targeting 3–6-year-olds had no significant effect on maternal employment. We interpret our results as evidence of a direct effect of childcare provision on maternal employment. We exclude that demand factors are driving these results because as we have shown in Table 1, the creation of workplaces in the childcare sector seems quite small. In addition, as a test, we estimated the effect of the childcare reform on mothers whose youngest child is aged 3–6 years, where we would not expect any effect shortly after the reform. This is confirmed by the empirical results (results available upon request).

A question could be raised to as how the quantitative effects of our study compare with previous work. We find that a 10 percentage point increase in childcare increases employment by 2.6–3.2 percentage points in the short term, that is, the first 3 years after the reform. This is the range of estimates reported in Tables 3 and 4 when we use the total target group and sub-groups of mothers. To facilitate interpretation, this translates into a 4.1–5.1 per cent increase in baseline maternal employment. In the long term, that is, 3–5 years after the reform, we find an effect of between 2.5–4.6 percentage points, or 4.0–7.3 per cent of baseline maternal employment (63 per cent). Most comparable in terms of the type of childcare reform are studies analysing universal reforms, as in Baker et al. (2008) in Canada, Bauernschuster and Schlotter (2015) in Germany and Nollenberger et al. (2015) in Spain. Given that the situation in Canada, and in Quebec, the region of the reform in 1997, is quite comparable to Norway before its reform in terms of childcare coverage and maternal employment, we may validly compare the corresponding qualitative and quantitative results. The situation in Germany and Spain, however, is quite different, with only a low provision of childcare and low maternal employment in long hours and full-time jobs in those countries.

In detail, Baker et al. (2008) estimate a differences-in-differences model and analyse the Quebec reform targeting children aged up to 4 years old[36]. Since they control for fixed effects in each province and for each year, the effect of the childcare policy in Quebec is identified by the change in Quebec, relative to other provinces, in 2000 or later relative to 1997 or earlier.

[36] In 1997, the government of Quebec introduced a subsidized universal childcare program to provide regulated childcare spaces to all children aged 0–4 years in Quebec at a parental contribution of C$5.00 per day. Children were eligible regardless of whether the parents were working. The program was phased in, starting with 4-year-olds in September 1997. Subsequently, 3-year-olds became eligible in September 1998, 2-year-olds in 1999, and children aged 0 and 1 in September 2000.

They estimate the ITT effect, as they estimate the reduced-form effect on all children, not only those using childcare. They estimate a 7.7 percentage point increase in employment in Quebec relative to the rest of Canada, which amounts to 14.5 per cent of baseline participation (57 per cent). This effect is for 4 years after the reform, which is close to what we consider the short-term effect.

However, unlike our study, Baker et al. (2008) define the reform variable as a dummy variable, whereas we estimate a linear effect in the pre-reform coverage rate. Apart from this, the childcare situation in Quebec before the reform is quite similar to that in Norway. Before the reform in 1997 in Quebec, 41 per cent of children were in childcare and 53 per cent of mothers were employed; after the reform, 62 per cent of children were in childcare and 63 per cent of mothers were employed. Hence, in our analysis, a 40 percentage point increase in childcare coverage produces similarly high effects when we use the entire target group, and a 20 percentage point increase in childcare coverage produces similar effects for mothers with more than one child and not on maternity leave. The Norwegian reform we study increased the percentage of children aged 1–2 years in childcare from 40 to 80 per cent between 2002 and 2008.

Baker et al. (2008) also show that crowding out or transitions from informal to formal care is taking place, but also that some enter childcare without the entry into work of the mother. We show for Norway that before the reform, 40 per cent of children were enrolled in public childcare, but that 72 per cent of mothers with a youngest child aged 2 years or less were registered as employed. This suggests that informal care during ordinary working hours also played an important role before the reform, as well as possibly other flexible work arrangements and not-ordinary work[37] that enabled parents to jointly cover longer childcare hours. Baker et al. (2008) demonstrate a decline in informal childcare, in the parental or other home, through the Quebec reform.

8 Conclusion

This study provides supportive quantitative evidence that making childcare universally available for 1–2-year-olds leads to new entries of mothers into work, increased maternal employment overall and increases in the actual hours of work. We find for all mothers whose youngest child is 2 years or younger, strong evidence of new entries into work as a result of the

[37] Not ordinary work is work outside the ordinary working hours, 6am to 6pm.

reform. We also find positive effects at the intensive margin only for mothers after maternity leave, the period following the first year after childbirth, that is, they transition into longer actual hours of work. However, actual hours are adjusted more within the range of hours less than full-time hours. The effects are mostly initiated shortly after the reform. However, the sub-group of mothers with more than one child are also more responsive in the long term following the reform.

Effects on labour supply will arguably lead to increased individual income and tax payments by this group. Tax revenue by the government is crucial to finance public childcare. The positive labour supply effects on both the extensive and intensive margins indicate that limited access to childcare is a constraint to parental, or more particularly, maternal employment. Hence, if childcare is not sufficiently available, government tax revenue will be lower because some mothers who are willing to work cannot because of their high opportunity cost. Overall, it seems surprising that responses in actual hours are not stronger and that we do not find evidence of the increased incidence of full-time work. This result may reveal that the institutional setting of maternity leave is likely to work as a disincentive to adjusting actual hours during maternity leave, along with contracted hours. Only after the job-protected leave do we find that mothers adjust hours. Other constraints or frictions may also exist that inhibit the ability of mothers to adjust hours upward. The time effects of the reform may also reflect a phase-in or short-term effects of the reform. It is noteworthy that the chances are increasing to 50 per cent and above to get a childcare place. This suggests that parents need high security in potential access to childcare to plan their labour market careers. This would be interesting to investigate further in other research.

These results speak to the policy debate about what the level of full childcare coverage should be to meet the demand by mothers who work or want to work. We find for a country with already high pre-reform female employment still remarkable positive effects on female employment through the expansion of childcare coverage from 40 to 80 per cent. In other words, productive female human capital in the market is increased through universal childcare reform. This demonstrates extra potential for other countries with low childcare coverage.

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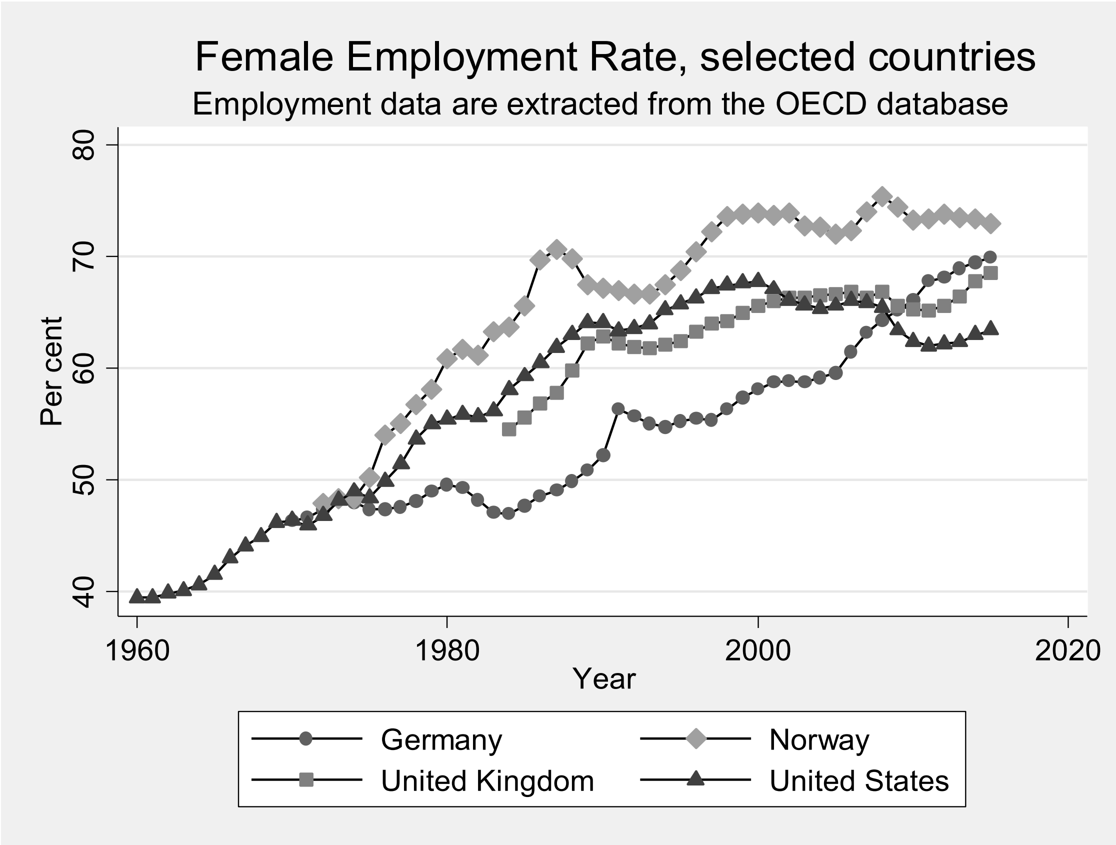
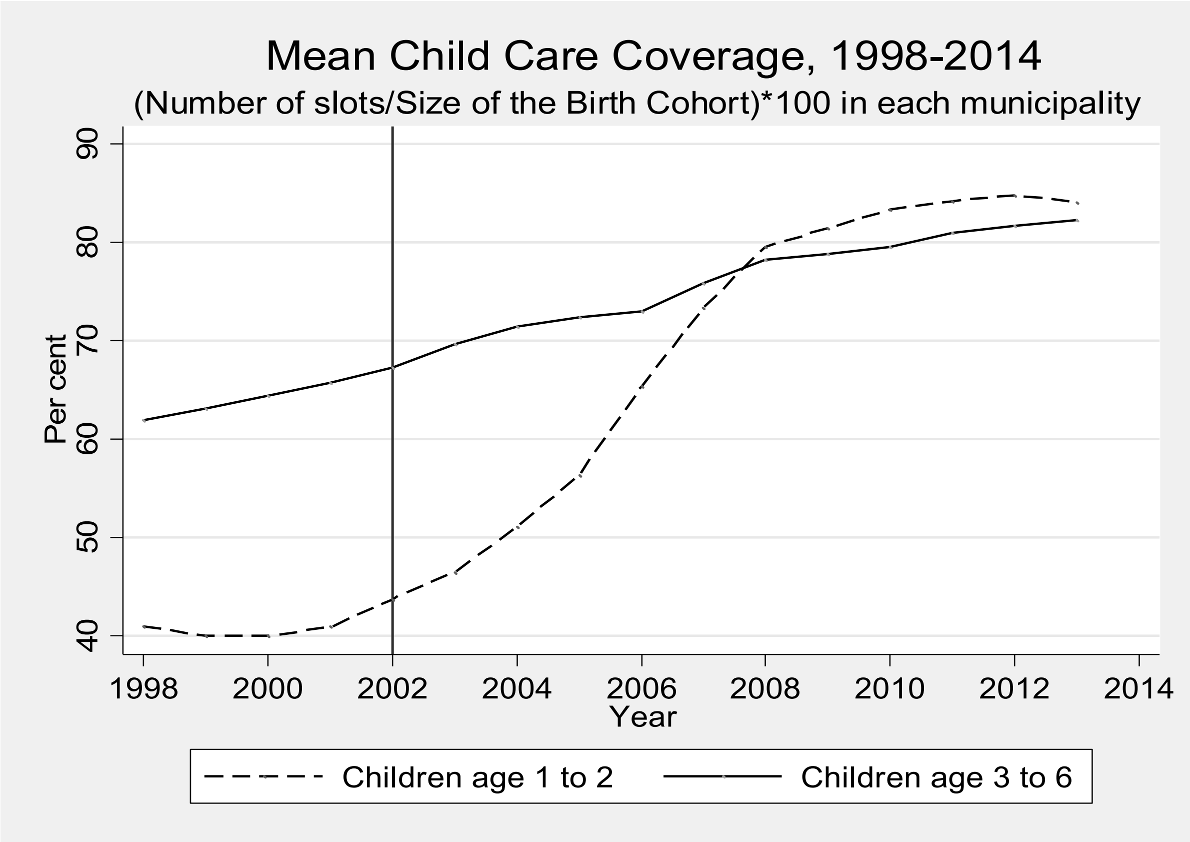
Figure 1



Female Employment Rate, selected countries
Employment data are extracted from the OECD database
Per cent
40
50
60
70
80
1960
1980
2000
2020
Year
Germany
Norway
United Kingdom
United States

Figure 2



Mean Child Care Coverage, 1998-2014
(Number of slots/Size of the Birth Cohort)*100 in each municipality
Per cent
40
50
60
70
80
90
1998
2000
2002
2004
2006
2008
2010
2012
2014
Year
Children age 1 to 2
Children age 3 to 6

Figure 3: Geographical distribution of child care across municipalities, pre-reform

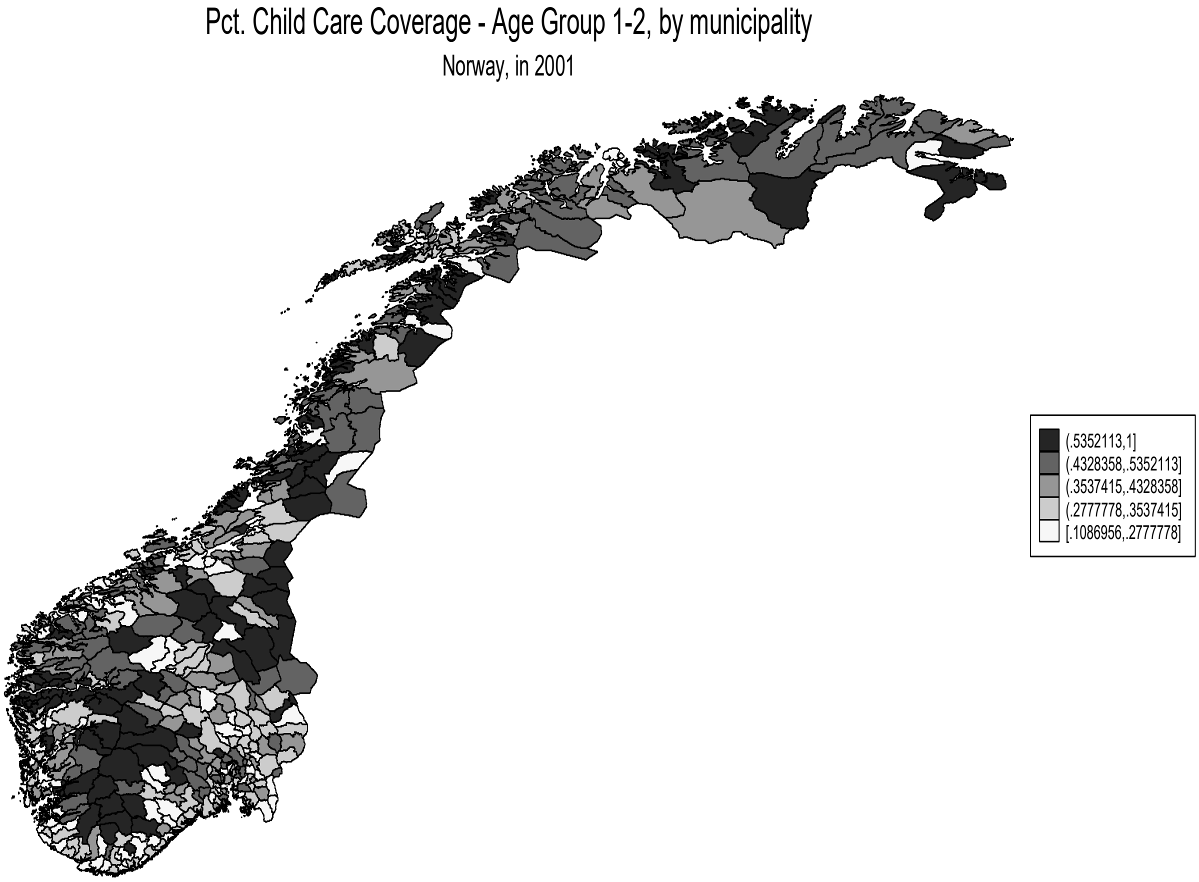


Figure 4.1

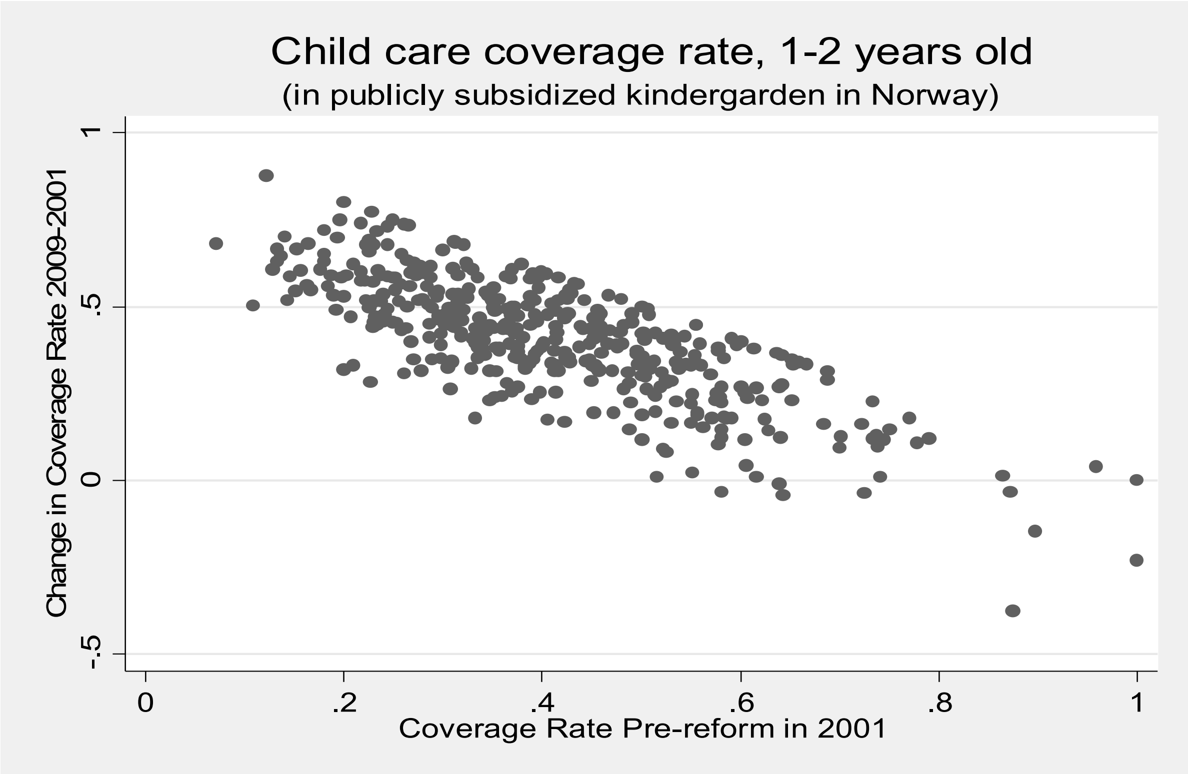


Figure 4.2

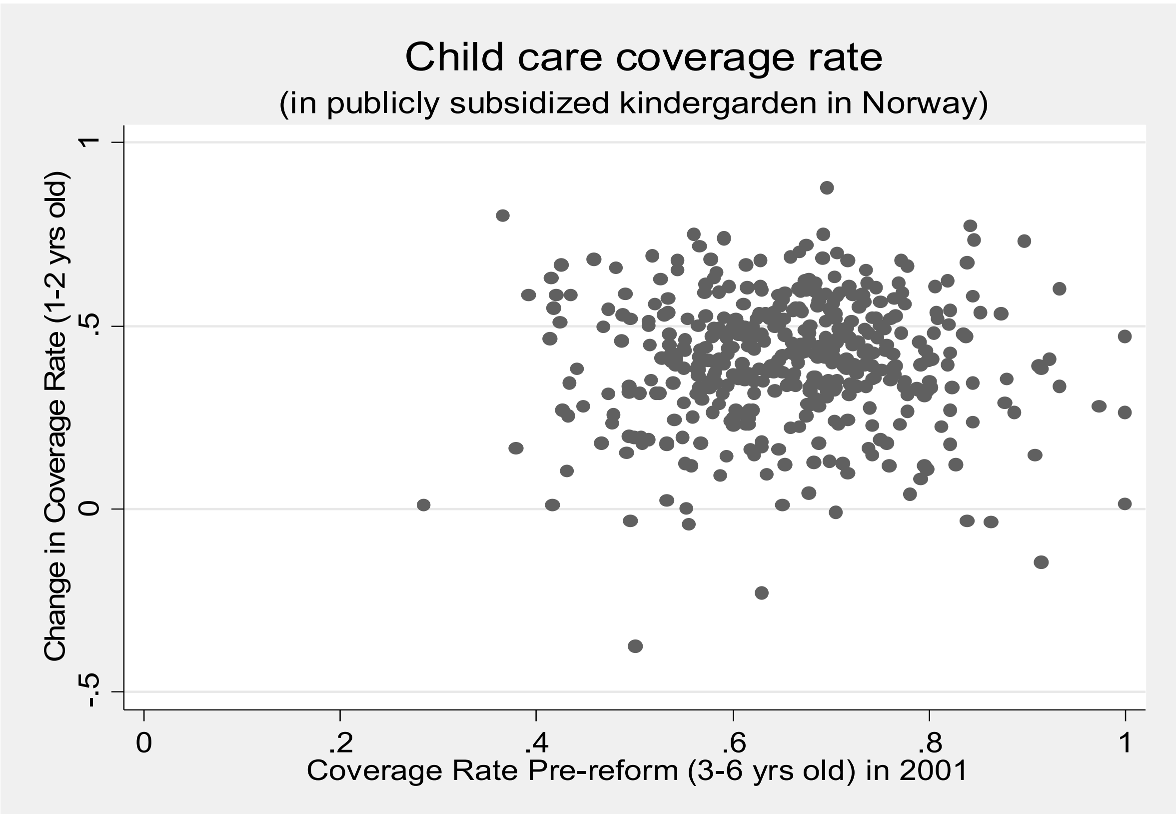


Table 1
Female employment in the kindergarten sector, 1980 - 2010

Year	**Number of employees in kindergarten sector (Male in brackets)**	**Percentage change over five years**	**Percentage of women working in kindergarten sector**
1990	35,891(900)	56%	3.8%
1995	51,832	44.4%	5.4%
2000	52,673	1.6%	4.9%
2005	64,713(6202*)	22.8%	5.4%
2010	87,401(9002)	35%	6.6%

Source: NOU 2012/1. Statistics Norway (www.ssb.no). * in 2006.

Table 2
Summary statistics for the pre-reform period 2001 ∗

Variables	Total	Mothers without maternity leave	Mothers with more than one child	Mothers with more than one child and without maternity leave
		Outcomes		
Work	0.72	0.63	0.72	0.62
	(0.45)	(0.48)	(0.45)	(0.48)
Long hours ∗∗	0.55	0.46	0.53	0.44
	(0.50)	(0.50)	(0.50)	(0.50)
Actual hours of work ∗ ∗ ∗∗	11.29	15.23	10.85	14.61
	(15.52)	(16.28)	(15.19)	(15.99)
Contracted hours of work	22.08	18.25	20.98	17.26
	(15.95)	(16.21)	(15.68)	(15.82)
		Distribution of hours ∗ ∗ ∗ ∗ ∗		
$0 < h \leq 18.5$	12.37	12.33	12.56	12.49
$18.5 < h \leq 30$	24.94	24.84	24.62	24.39
$30 < h \leq 37.5$	36.88	36.79	36.80	36.70
$37.5 < h$	39.57	39.67	39.66	40.04
$h = 0$	0.28	0.37	0.28	0.38
		Individual controls		
Age	31.26	31.33	32.87	33.07
	(5.05)	(5.27)	(4.56)	(4.72)
Education in years	14.15	14.01	14.03	13.89
	(2.10)	(2.11)	(2.10)	(2.08)
Number of children	1.88	1.91	2.46	2.50
	(0.89)	(0.93)	(0.69)	(0.73)
Married ∗ ∗ ∗	0.93	0.91	0.95	0.93
	(0.26)	(0.28)	(0.22)	(0.25)
N	3,439	2,549	2,076	1,542

Notes:
∗ The main sample of mothers contains mothers with the youngest child age 2 or younger observed in 2001.
∗∗ Long hours refers to weekly working hours 18.5 and above.
∗ ∗ ∗ Marital status is: 1=married; 0=single/divorced/widowed.
∗ ∗ ∗∗ Hours of work include zero hours.
∗ ∗ ∗ ∗ ∗ Group averages are presented for the distribution of contracted hours.

Table 3
Main regression results: All mothers with the youngest child age 2 and below

	Employment	Long hours	Actual hours	Contracted hours
Short-term ITT 2002-2004	0.247**	-0.045	3.992	10.434*
	(0.104)	(0.115)	(9.335)	(5.749)
Long-term ITT 2005-2009	0.203*	0.021	6.330	7.223
	(0.119)	(0.138)	(8.987)	(6.131)
Pre-reform mean	0.72	0.55	11.30	22.08
Number of Observations	27,577	27,577	27,577	27,577
Individual characteristics 1)	Yes	Yes	Yes	Yes
Time dummies	Yes	Yes	Yes	Yes
Municipality fixed effects	Yes	Yes	Yes	Yes

Note:
The coefficients reported in this table are based on the interaction term between the pre-reform child care coverage rate in 2001 and the time dummy. The reported coefficients are multiplied by -1 to reflect the effects of the policy change, because lower coverage rate implies higher treatment intensity. The complete results are in the Appendix Table A. Standard errors are in parentheses and are clustered at the municipal level * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$
1) Control variables for the individual characteristics include age, age squared, years of education, number of dependent children, and dummy variables for marital status.

Table 4
Regression results: sub-groups of mothers with the youngest child age 2 and below

	Employment	Long hours	Actual hours	Contracted hours	Number of Observations
Without Maternity Leave					
Short-term ITT	0.265**	-0.115	12.808*	12.581*	20,435
2002-2004	(0.124)	(0.138)	(7.519)	(6.945)	
Long-Term ITT	0.254*	0.009	11.334	12.215	
2005-2009	(0.143)	(0.167)	(7.844)	(7.687)	
Dependent Mean	0.63	0.46	15.23	18.25	
With More than One Child					
Short-term ITT	0.269*	0.038	11.172	11.435	16,873
2002-2004	(0.139)	(0.166)	(10.942)	(7.604)	
Long-Term ITT	0.357**	0.157	19.952*	13.634*	
2005-2009	(0.138)	(0.179)	(12.032)	(7.328)	
Dependent Mean	0.72	0.53	10.85	20.98	
With More than One Child and No Mat Leave					
Short-term ITT	0.317*	0.025	17.053*	13.683	12,481
2002-2004	(0.174)	(0.188)	(9.409)	(9.476)	
Long-Term ITT	0.461***	0.236	22.260**	19.867**	
2005-2009	(0.175)	(0.214)	(10.228)	(9.366)	
Dependent Mean	0.62	0.44	14.61	17.26	
Individual characteristics 1)	Yes	Yes	Yes	Yes	
Time dummies	Yes	Yes	Yes	Yes	
Municipality fixed effects	Yes	Yes	Yes	Yes	

Note:
The coefficients reported in this table are based on the interaction term between the pre-reform child care coverage rate in 2001 and the time dummy. The reported coefficients are multiplied by -1 to reflect the effects of the policy change, because lower coverage rate implies higher treatment intensity. The complete results are avaliable on request. Standard errors are in parentheses and are clustered at the municipal level * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$
1) Control variables for the individual characteristics include age, age squared, years of education, number of dependent children, and dummy variables for marital status.

Table 5
Treatment effects on fertility

	Newborn	Number of Observations
All municipalities		
Short-term ITT	0.012	249,716
2002-2004	(0.013)	
Long-term ITT	0.013	
2005-2009	(0.013)	
Individual characteristics 1)	Yes	
Time dummies	Yes	
Municipality fixed effects	Yes	

Notes:
The dependent variable is the incidence of a new born child at the individual mother level from the LFS for the years 2001-2009. The variable newborn is a dummy variable equal to 1 if there was a child born to a woman in a given year.
The coefficients reported are based on the interaction term between child care coverage rate in 2001 and the time dummy. The reported coefficients are multiplied by -1 to reflect the effects of the policy change, because lower coverage rate implies higher treatment intensity. The complete results are avaliable on request. Standard errors are in parentheses and are clustered at the municipal level
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$
1) Control variables for the individual characteristics include age, age squared, years of education, number of dependent children, and dummy variables for marital status.

Table 6
Regression results: PLACEBO test

	Employment	Long hours	Actual hours	Contracted hours	Number of Observations
			All		
Short-term ITT 2000	-0.006	-0.001	-1.030	4.549	14,432
	(0.119)	(0.125)	(9.906)	(5.543)	
Long-Term ITT 2001	-0.077	-0.019	2.894	1.146	
	(0.141)	(0.169)	(10.895)	(6.895)	
			Without Maternity Leave		
Short-term ITT 2000	-0.029	0.088	-0.371	1.970	10,656
	(0.143)	(0.160)	(8.394)	(7.607)	
Long-Term ITT 2001	-0.097	0.054	-1.444	-2.091	
	(0.185)	(0.218)	(9.999)	(9.690)	
			With More than One Child		
Short-term ITT 2000	0.105	0.022	6.350	9.343	8,600
	(0.165)	(0.174)	(12.423)	(7.259)	
Long-Term ITT 2001	-0.059	-0.108	0.601	-1.241	
	(0.173)	(0.222)	(14.440)	(8.478)	
			With More than One Child and No Mat Leave		
Short-term ITT 2000	0.147	0.147	4.561	9.284	6,388
	(0.197)	(0.218)	(11.219)	(9.889)	
Long-Term ITT 2001	-0.063	-0.025	-1.933	-3.201	
	(0.231)	(0.287)	(12.830)	(12.142)	
Individual characteristics 1)	Yes	Yes	Yes	Yes	
Time dummies	Yes	Yes	Yes	Yes	
Municipality fixed effects	Yes	Yes	Yes	Yes	

Note:
This table reports placebo results for the overall sample and the three sub-samples using the period 1998-2001. The placebo reform is set to 2000 in this test.
The coefficients reported in this table are based on the interaction term between the pre-reform child care coverage rate in 1998 and the time dummy. The reported coefficients are multiplied by -1 to reflect the effects of the policy change, because lower coverage rate implies higher treatment intensity. The complete results are avaliable on request. Standard errors are in parentheses and are clustered at the municipality level * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.
1) Control variables for the individual characteristics include age, age squared, years of education, number of dependent children, and dummy variables for marital status.

Appendix Table A1: Selected studies on maternal labor supply responses to child care expansions targeting children age 3 and younger								
	Country	Policy/Year	Sample	Year	Method	Target group	Causal effect on Participation/employment rate (percentage point) (st. errors)	Causal effect on Hours, weekly (st. errors)
Bauernschuster and Schlotter (2015)	Germany	Introduction of a legal claim to a place in kindergarten in 1996	Socioeconomic panel (SOEP) and Micro-census	1992-2000	DID	Mothers with youngest child age 3-4	3.4***	2.5, own calculation based on table2 and table4
Bettendorf, Jongen, and Muller (2015)	Netherlands	Increased Subsidy to Child care 2005	Dutch Labor Force Survey	1995-2009	DID	Mothers with child age 3 and below	2005-2007: 1.8***(0.007) 2008-2009: 2.5***(0.007)	2005-2007: 1.009***(0.218) 2008-2009: 1.418***(0.215)
Nollenberger and Rodriguez-Planas (2015)	Spain	Expansion in child care 1991, 1993, 1995 and 1996,	Spanish Labor Force Survey	1987-1997	DDD	Mothers with youngest child 3 years old	Average across multiple implementation years: 2.8*(1.6)	NA
Lefebvre, Merrigan, and Roy-Desrosiers (2012)	Quebec, Canada	Staged introduction of low-cost child care starting from 1997	7 bi-annual cycles of National Longitudinal Survey of Children and Youth	1994-2006	DID	Mothers with child age 1-4	8.0%***(2.0) in 1999 – 13.0%***(2.0) in 2007	NA means there is no results on hours
Lefebvre and Merrigan (2008)	Quebec, Canada	Staged introduction of low-cost child care starting from 1997	Survey of Labor and Income Dynamics	1993-2002	DID	Mothers with child age 1-5	7.3***(2.6)	0.31 hours/week, calculated based on estimated effects of annual hours (133***(52)) and annual weeks (4.28***(1.33))
Baker, Gruber, and Milligan (2008)	Quebec, Canada	Staged introduction of low-cost child care starting from 1997	4 bi-annual cycles of National Longitudinal Survey of Children and Youth	1994-2003, excluding 1998/1999 cycle	DID	Mothers with child age 0-4	7.7***(0.6)	NA
Goux, and Maurin (2010)	France	Child age dependent public preschool attendance rule 1999	French population full census	March 1999	RDD and IV	Single mothers with youngest child age 2 or 3	Single mothers of child age 2: -1.1(1.4) Single mothers of child age 3: 8.6***(1.0)	NA

Table A2
Complete results: mothers with the youngest child age 2 and below

	Employment	Long hours	Actual hours	Contracted hours
Short-term ITT	-0.247**	0.045	-3.992	-10.434*
2002-2004	(0.104)	(0.115)	(9.335)	(5.749)
Long-term ITT	-0.203*	-0.021	-6.330	-7.223
2005-2009	(0.119)	(0.138)	(8.987)	(6.131)
Post-short	0.082**	-0.030	0.064	3.669
2002-2004	(0.043)	(0.047)	(3.794)	(2.244)
Post-long	0.113**	0.011	4.338	5.100***
2005-2009	(0.050)	(0.054)	(3.657)	(2.441)
Age	0.089***	0.075***	4.537***	5.099***
	(0.009)	(0.011)	(0.576)	(0.502)
Age2	-0.001***	-0.001***	-0.054***	-0.070***
	(0.000)	(0.000)	(0.009)	(0.008)
Years of education	0.031***	0.030***	1.269***	1.677***
	(0.002)	(0.002)	(0.140)	(0.115)
Number of children	-0.064***	-0.078***	-4.090***	-4.376***
($< 16yrs$)	(0.009)	(0.009)	(0.471)	(0.418)
Single	-0.216***	-0.149***	-6.166***	-10.068***
	(0.017)	(0.015)	(1.342)	(0.901)
Widowed/Divorced	-0.241***	-0.166***	-7.169**	-10.933***
	(0.047)	(0.043)	(3.410)	(2.552)
Constant	-1.151	-1.072	-99.958	-87.425
	(0.136)	(0.158)	(9.088)	(7.558)
Adjusted R^2	0.141	0.109	-	-
Number of Clusters	404	404	404	404
Number of Observations	27,577	27,577	27,577	27,577
Municipality FE	Yes	Yes	Yes	Yes

Note:
Standard errors are in parentheses and are clustered at the municipal level * p < 0.10, ** p < 0.05, *** p < 0.01.

Table A3
Complete results: mothers without maternity leave

	Employment	Long hours	Actual hours	Contracted hours
Short-term ITT	-0.265**	0.115	-12.808	-12.581*
2002-2004	(0.124)	(0.138)	(7.519)	(6.945)
Long-term ITT	-0.254*	0.009	-11.334	-12.215
2005-2009	(0.1430)	(0.167)	(7.844)	(7.687)
Post-short	0.085**	-0.067	3.472	4.426
2002-2004	(0.051)	(0.056)	(3.043)	(2.746)
Post-long	0.148**	0.028	5.930*	8.184***
2005-2009	(0.060)	(0.064)	(3.226)	(3.148)
Age	0.098***	0.073***	5.595***	6.207***
	(0.009)	(0.011)	(0.590)	(0.572)
Age2	-0.001***	-0.001***	-0.074***	-0.082***
	(0.000)	(0.000)	(0.009)	(0.009)
Years of education	0.034***	0.029***	1.876***	1.955***
	(0.002)	(0.003)	(0.150)	(0.147)
Number of children	-0.074***	-0.072***	-4.412***	-5.146***
($< 16yrs$)	(0.009)	(0.010)	(0.595)	(0.531)
Single	-0.190***	-0.114***	-9.961***	-9.970***
	(0.019)	(0.016)	(1.167)	(1.109)
Widowed/Divorced	-0.212***	-0.139***	-11.636***	-10.660***
	(0.049)	(0.044)	(3.082)	(3.109)
Constant	-1.449	-1.144	-112.771	-117.267
	(0.136)	(0.163)	(9.243)	(8.790)
Adjusted R^2	0.161	0.112	-	-
Number of Clusters	395	395	395	395
Number of Observations	20,435	20,435	20,435	20,435
Municipality FE	Yes	Yes	Yes	Yes

Note:
Standard errors are in parentheses and are clustered at the municipal level * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table A4
Complete results: mothers with more than one child

	Employment	Long hours	Actual hours	Contracted hours
Short-term ITT	-0.269**	-0.038	-11.172	-11.435
2002-2004	(0.139)	(0.166)	(10.942)	(7.604)
Long-term ITT	-0.357**	-0.157	-19.952*	-13.634*
2005-2009	(0.138)	(0.179)	(12.032)	(7.328)
Post-short	0.091	0.005	3.672	4.293
2002-2004	(0.056)	(0.064)	(4.525)	(2.889)
Post-long	0.169***	0.062	10.614**	7.816***
2005-2009	(0.060)	(0.072)	(4.822)	(2.997)
Age	0.103***	0.092***	5.665***	5.842***
	(0.015)	(0.016)	(1.065)	(0.790)
Age2	-0.001***	-0.001***	-0.069***	-0.080***
	(0.000)	(0.000)	(0.015)	(0.011)
Years of education	0.033***	0.030***	1.408***	1.832***
	(0.004)	(0.004)	(0.227)	(0.167)
Number of children	-0.088***	-0.071***	-5.086***	-5.434***
($< 16yrs$)	(0.012)	(0.012)	(0.837)	(0.563)
Single	-0.209***	-0.154***	-4.316**	-9.604***
	(0.030)	(0.023)	(1.788)	(1.402)
Widowed/Divorced	-0.252***	-0.163***	-8.767***	-11.456***
	(0.058)	(0.049)	(4.130)	(3.283)
Constant	-1.346	-1.379	-122.083	-99.101
	(0.260)	(0.285)	(18.654)	(13.764)
Adjusted R^2	0.151	0.111	-	-
Number of Clusters	385	385	385	385
Number of Observations	16,783	16,783	16,783	16,783
Municipality FE	Yes	Yes	Yes	Yes

Note:
Standard errors are in parentheses and are clustered at the municipal level * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table A5
Complete results: mothers with more than one child and without maternity leave

	Employment	Long hours	Actual hours	Contracted hours
Short-term ITT	-0.317*	-0.025	-17.053*	-13.683
2002-2004	(0.174)	(0.188)	(9.409)	(9.476)
Long-term ITT	-0.461***	-0.235	-22.260**	-19.867*
2005-2009	(0.175)	(0.214)	(10.228)	(9.366)
Post-short	0.111	-0.008	5.606	5.468
2002-2004	(0.071)	(0.075)	(3.950)	(3.731)
Post-long	0.231***	0.096	10.791**	11.667***
2005-2009	(0.075)	(0.084)	(4.256)	(3.913)
Age	0.115***	0.093***	7.046***	7.279***
	(0.017)	(0.017)	(1.017)	(0.971)
Age2	-0.002***	-0.001***	-0.094***	-0.097***
	(0.000)	(0.000)	(0.014)	(0.014)
Years of education	0.036***	0.028***	2.005***	2.115***
	(0.004)	(0.004)	(0.238)	(0.220)
Number of children	-0.096***	-0.067***	-5.677***	-6.192***
($< 16yrs$)	(0.012)	(0.013)	(0.794)	(0.722)
Single	-0.174***	-0.105***	-9.401**	-8.789***
	(0.028)	(0.023)	(1.671)	(1.558)
Widowed/Divorced	-0.211***	-0.134***	-12.142***	-10.692***
	(0.059)	(0.048)	(3.766)	(3.913)
Constant	-1.746	-1.526	-137.640	-135.549
	(0.297)	(0.315)	(18.111)	(17.341)
Adjusted R^2	0.174	0.118	-	-
Number of Clusters	376	376	376	376
Number of Observations	12,481	12,481	12,481	12,481
Municipality FE	Yes	Yes	Yes	Yes

Note:
Standard errors are in parentheses and are clustered at the municipal level * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

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