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The effect of fertility timing on labor market work duration

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The effect of fertility timing on labor market work duration[a]

by

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Abstract
We provide a framework for the estimation of the impact of fertility timing on female long-term labor supply, measured as labor market work duration. We show that the genuine treatment is waiting time to birth rather than birth per se. In the application we control for the joint decision of fertility and labor supply by using the 'same-sex' instrument in a control function setting. We find that having a third child will in general reduce the labor market work duration. The magnitude of the effect depends to a large extent on the mothers' age at second birth but also on the waiting time to the third child and the education level.

Keywords: Labor supply; Parenthood; Retirement; Dynamic treatment assignment; Censored regression
JEL-codes: J22; C24; C26; C51

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1 Introduction

One of the most prominent labor market phenomena of the 20th century is the dramatic increase in the labor supply of married women, and the parallel decline in fertility rates in OECD countries. The general findings from quasi-experimental approaches (e.g. twins adopted by Rosenzweig and Wolpin (1980) and Bronars and Grogger 1994; and the same-sex sibling approach introduced by Angrist and Evans 1998) is that labor income is reduced as a consequence of having a child. These findings seem to suggest that high fertility rates are difficult to combine with a high female labor supply. Yet while these empirical findings seem to be robust, they fail to explain why there are countries where both the female labor supply and the fertility rates are high (e.g. the Scandinavian countries). In order to understand these seemingly conflicting results, it is important to understand the dynamics of the labor supply and how it is related to fertility.

The present paper provides a methodological framework to capture the effect of fertility on long-term labor supply, measured as labor market work duration, and applies it to Swedish data. Although the framework does not provide direct estimates of the effect of having a child on life-time income, the effect on work duration can be interpreted as a lower bound, in absolute terms, of the effect on lifetime income. This is because life-time income is a function of work duration and wages, and the effect of fertility on wages can be assumed to be negative, if any.

The methodological framework shows that the genuine treatment variable is waiting time to fertility rather than fertility per se. For example, there is a waiting time to: (i) motherhood from birth, (ii) motherhood from the onset of working life for women who start to work before becoming mothers, (iii) a second child for mothers with one child, (iv) a third child for mothers with two children, etc. In all these situations, the waiting times are chosen. An inherent characteristic of the waiting time to children is that it is observed for women with completed fertility, while it is censored for fertile women at the sampling date (i.e., end of the study).

The identification of the effect of fertility is difficult as the timing of pregnancy and career decisions most likely are jointly determined. This is a problem that, most likely, is especially severe if we aim at estimating the effects for the population of mothers (i.e. (i)), and to a somewhat lesser degree, for populations (ii)-(iv). In order to estimate a causal effect for any of these populations some form of exogenous shock is warranted. In this paper we focus on (iv) and use the fact that parents have preferences for a mixed sex sibling composition which exogenously increases the likelihood of a third child for mothers with two children of the same sex (cf. Angrist and Evans 1998).

Previous studies using the ‘same-sex’ instrument are based on cross-sectional data on income or wages at the sampling date. Any estimator using income data based on this sampling strategy provides a weighted average of potential effects of a third child on income at different durations after the third child is born. Assume no effect on leaving the labor market early but a short term reduction of working time over two or three years. Then this estimator will show a labor supply reduction even though it has no permanent effect on leaving the labor market.[1] The result would provide evidence of a cost of parenthood, but as this cost is shared equally for all mothers, the cost on life time income for each single mother will be quite low. This is especially true if a maternity leave insurance with income replacement, as is the case in Sweden. If however the majority of women return to work fast but some women are gradually leaving the labor market the effect of motherhood on life time income would be large for a few of these women. Being able to distinguish between long and short term consequences and to identify effect heterogeneity is hence important in the design of any social insurance system. The Swedish social insurance system has income replacement after parenthood based on previous earnings in combination with strict employment protection regulations with regard to pregnancy and child rearing. Such a system would theoretically give small effects on long-term labor supply and at the same time provide high fertility rates. Empirically, this is an open question and our understanding is that ours is the first paper evaluating the consequences of motherhood on long-term labor supply measured as labor market work duration.

As the same sex instrument affects fertility (or rather the hazard rate of fertility) it also

[1] For example, Angrist and Evans (1998) find that having a third child reduces participation by 9–10 % points; Hyslop (1999) estimates that a child 0–2 years old reduces participation by 11–17 % points; and Rosenzweig and Wolpin (1980) find that the second child reduces the participation probability by 37 % points.

affects the spacing between the second and third child. There is some evidence that birth spacing affects female labor supply.[2] Troske and Voicu (2012) show that both the timing and the spacing of births matter for women's labor supply around birth. Karimi (2014) uses miscarriages between the first and second child as an instrument for the change in child spacing. She finds that a one-year delay of second births increases the probability of re-entering the labor market between births. In addition, a one-year delay increases labor income (up to 15 years after the second-born child). The bottom line is that waiting may be an important aspect to acknowledge in any analysis of the effect of fertility on female labor supply.[3]

In the estimation we use data from Swedish administrative registers, providing total coverage for those who were born in 1923 or later. We sample mothers to second-born children from the register and then – unless not censored due to end of study – we can follow them until retirement. We restrict the analysis to mothers born in between 1923 and 1947. The reason is that we do not want to have too much censoring. With this choice the fraction of right- and left-censored is 16 percent. The final analysis data consist of 804,721 mothers. However, as there were large changes in the social insurance system in the 1970's we also restrict the analysis to 374,932 mothers born between 1923 and 1935. That is, we use two sets of data: one for mothers born in 1923-1947, and the other for 1923-1935.

The results from the estimation (a double-censored MLE with a control function) is that having a third child will in general reduce the labor market work duration for mothers with three children. The magnitude of the effect depends to a large extent on the age of the mother when giving second birth, but also on the waiting time to the third child after the second, and the education level. For mothers with three children with average education, average age at second birth, and the shortest spacing between second and third birth, the

[2]Child spacing has also been seen to be affected by polices. Lalive and Zweimüller (2009) show e.g. that the Austrian parental leave system affected both employment of mothers, the number of children, and the spacing of births. Hoem (1993) shows that parents decreased the spacing as consequence of an administrative rule in the Swedish parental leave system where one could retain the same level of benefits for the subsequent child without having to return to work between births.

[3]Increased child spacing has not been seen to affect women's labor supply. It has also been seen as a means of improving both maternal and infant health (see e.g. Rosenzweig 1986 and Buckles and Munnich 2012).

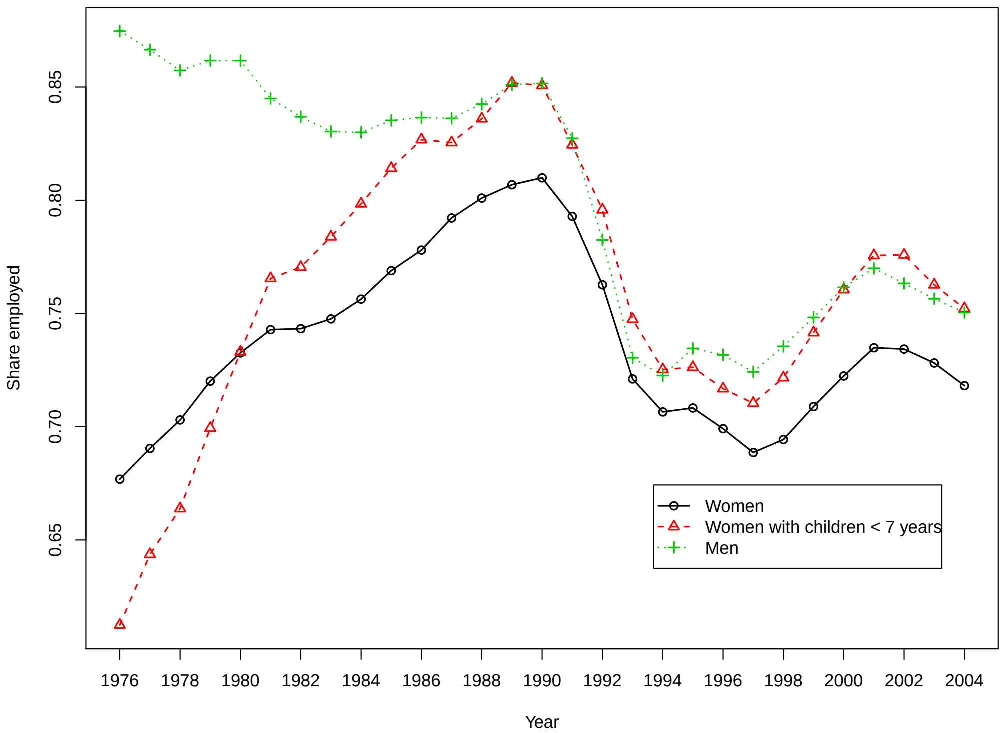

Figure 1: Labor force participation among all Swedish men, women, and women with children under 7 years between 1976 and 2004. Source: AKU, Statistics Sweden.

reduction in the labor market duration is 5.2 percent. For highly educated mothers who had their second child early, the reduction in labor market work duration can be even more than 30 percent. For low-educated mothers with a second child born late and short second-to-third child spacing, we even find statistically and economically significant positive effects of having a third child on labor market work duration. The results from the more restrictive sample are qualitatively the same, however the effects are somewhat smaller in magnitude (e.g. an average effect of 3.2 percent for mothers with short child spacing). Given that the changes in the tax and social insurance system in the 1970's increased women's labor supply this result could be seen as a surprise. However, as we estimate effects for mothers with three children and not effects of parenthood per se the reduction in magnitude may be less surprising. Two-child mothers born early are probably a more positively selected sample of women with better labor market attachment than later cohorts.

In the next section we show how the female labor supply in Sweden has developed since the mid 1970s and we also provide a short presentation of the Swedish institutions. Section 3 provides the framework for evaluation, Section 4 describes the data. The results are presented in section 5 and finally, section 6 concludes.

2 Female labor supply and the Swedish labor market

From an international perspective, the Swedish labor market stands out with its high female labor supply (see e.g. Angelov et al. 2016). To describe the historical evolvement of these high rates in more detail, Figure 1 shows the share of working Swedish men, women, and women with children younger than seven years for the period 1976 to 2010. From this figure, we can see that the female labor market participation rate in Sweden increased rapidly and without interruption from 1976 until the early 1990s. After this period, the gender gap in labor market participation has remained essentially unchanged, and there has been a gradual increase in the participation rates of both men and women.

In the following we briefly present the Swedish institutional context and discuss possible reasons for the high labor supply among Swedish women. Several reforms, starting already in the 1930's, have contributed to the high labor supply among women. A maternity allowance was introduced in the early 1930's which provided the first statutory right to compensation in connection with childbirth (SFS 1937:338). In the early 1950's a regulation was taken which protected women against dismissal, firing and deteriorating employment conditions due to circumstances that could be attributed to pregnancy or birth (SOU 1946:60). From the 1950's there was a gradual increase in the coverage of the maternity allowance together with an increase in the benefits (Persson 2013). The introduction of the individual tax system in 1971, whereby taxation of spouses was individualized, created large incentives for Swedish women to participate in the labor force. Selin (2009) concludes that the female labor supply increased by 10 percentage points due to this reform.[4] Around the same time, in 1974, the parental leave system was introduced. The replacement rate for parental leave was from the very beginning proportional

[4] In this context it is interesting to note that, among the compared countries, Germany has the lowest female participation rate and in Germany couples are still taxed together.

to forgone earnings,[5] which probably has contributed to the high employment rate among women before entering parenthood. The generous replacement rate for parental leave and the flexibility of when to use the paid days probably also have contributed to the fact that most Swedish mothers labor market work while having small children.[6] In the present parantal leave system any one of the parents can stay at home on a full-time basis with job protection during the child's first 18 months. Parents can take turns being on parental leave, as long as the total number of months on leave is at most 18 months per child. Thereafter, parents are allowed to reduce their working hours up to 25 percent until the child turns 8 years old (SFS 1995:584). Women use the parental insurance most: they take out 80 percent of the paid parental leave days (Försäkringskassan, 2011). In addition, 44 percent of all women in the ages 25-54 work part-time (<35 hours per week). The corresponding share of men who work part-time is 10 percent.[7]

In parallel to the institutional changes described above, there has been a rapid increase in the public provision of child care, especially during the 1980s. In the first part of the 1980s only 30-35 percent of children below age 3 was enrolled in publicly subsidized care (Mörk et al., 2013). This may also have contributed to the high female labor supply, but it could also be a symptom of an increased demand for child care – a causal relationship between public provision of child care and female labor supply has not been established empirically.

The above mentioned reforms have coincided with a large increase in the labor force participation rate among Swedish women in general and mothers with young children in particular. As shown in Figure 1, the participation rate among women increased with 10 percentage points between 1976 and the climax of the economy boom in 1990. Among women with pre-school children, the increase was about twice as large: almost 20 percentage points. From 1990 onwards, the participation rate is higher among women with pre-school children, than among women in general. This is most likely a cohort effect,

[5]Forgone earnings is basically equal to previous earnings.

[6]The introduction of the Swedish parental leave system has not (to our knowledge) been evaluated with respect to its effect on female labor supply.

[7]Source: Public statistics from Statistics Sweden, published on the web: http://www.scb.se/Pages/Article_ _-_ _332715.aspx

that is, almost all young women participate in the labor force, while being a housewife is more common among older women.

3 Methodological framework

3.1 Notation and Variables

For a woman i with the second birth at calendar time point T_{2i}, let W_i^* be the waiting time until the 'potential' third birth time T_{3i}^*, and Y_i^* be the duration to the retirement time point T_{ri}^* from the calendar time point of having a second child:

$$W_i^* \equiv T_{3i}^* - T_{2i} \text{ and } Y_i^* \equiv T_{ri}^* - T_{2i} \Longleftrightarrow T_{ri}^* \equiv Y_i^* + T_{2i}$$

From these definitions, it follows that T_3^* is never observed for the women who do not have a third child until fertility ends at age s. In contrast, T_r^* and Y^* are realized always, although they may not be observed fully due to censoring problems, as we do not observe the whole labor market career of the women.[8] We will sometimes omit the subscript i under the i.i.d. assumption across $i = 1,...,N$.

In order to handle the intrinsic censoring of T_3^* we need to sample women who no longer can have children: we let $s = 45$. We have *income data* for the period 1985 to 2010 which we use to determine retirement status. Retirement is defined as having no income during c consecutive years. We set $c = 3$, but have performed sensitivity analyses with $c = 4$ and 5, showing the results not being sensitive to this choice. The mandatory retirement age was 65 until 1998, and given that the effective (average) retirement age was 62 we sample women born between 1923 and 1963.[9] Women born in 1923 were aged 45 in 1968 and were 65 years old in 1988 and women born in 1963 were 45 in 2008; we look at 1988 and 2008 instead of the income data period, 1985 and 2010, because of $c = 3$. The implication of this procedure is, thus, that both left and right-censored observations are

[8] If we want to know the effect on having a child, we would simply exchange T_{2i} with T_{li}, the time entering the labor market (the population is restricted to women entering the labor market pre birth) or T_{mi}, the time being married (the population is restricted to married couples without a child). W_i^* is then the duration/waiting time until first birth time, T_{1i}, $W_i^* \equiv T_{1i} - T_{li}$ or $W_i^* \equiv T_{1i} - T_{mi}$.

[9] The average retirement age has increased slowly over the study period. In 1980 the average was just above 61 and in 2010 the average was just above 62 (Olsson 2011).

possible.

To understand the left-censoring of Y^*, consider a woman born in 1923. Suppose her incomes over 1985-1987 are zero. Then she must have retired in 1985 (age 62) or earlier; i.e., her retirement age is equal to or smaller than 62, which implies a left-censoring of Y^* at $1985 - T_2$. The right-censoring occurs because there are many women who have not retired by 2010; for them, we only know $T_r^* \geq 2010$ that is equivalent to

$$Y^* \geq 2010 - T_2.$$

To understand the notation better, see the figure below:

Event:	born	2nd birth	3rd birth	fertility stop	retirement
Calendar					
Time Point:	$T_b \longrightarrow$	$T_2 \longrightarrow\longrightarrow$	$T_3^* \longrightarrow\longrightarrow$	$T_s \equiv T_b + s \longrightarrow$	$T_r^* \longrightarrow\longrightarrow$
Duration:		$W^* \equiv T_3^* - T_2$		$Y^* \equiv T_r^* - T_2 \geq 0$	

So long as the mother's birth year T_b is exogenous to the other variables, this way of selecting a sample based on T_b does not pose a sample selection problem; at worst, we may declare that our interest is on the birth cohorts in the above condition. We assume $T_s \leq T_r^*$ (no retirement before age s). This restriction is necessary because otherwise, T_r^* may precede the cause (having a third child). For instance, a woman may retire at age 30, which then affects her decision to have a third child at age 40. Since one can work anytime so long as alive, the definition of retirement is ambiguous anyway, and ruling out retirement before s is not too far-fetched.

Let $1[A] \equiv 1$ if A holds and 0 otherwise. Define the key dummy variable for having a third child or not:

$$\begin{aligned} D_i &\equiv 1[\text{woman } i \text{ has a third child before age } 45] \\ &= 1[T_{2i} + W_i^* - T_{bi} < 45] = 1[W_i^* < 45 - T_{2i} + T_{bi}]. \end{aligned}$$

Our interest is on how D and W^* affect the outcome variable Y^* for women with at least two children. A small W^* (i.e., having a third child soon after the second) may reduce the likelihood of returning to work between births, which means a longer spell of interrupted market work. This may lower the human capital, which can affect the labor supply by shortening the work duration Y^*. A large W^* may, on the other hand, also be non-optimal as women, by definition, deliver the child when older. There are, at least, two potential reasons why a delayed pregnancy can increase the hazard to retirement for women. An increased age of the mother may increase the risk of health problems for both the child and mother (see e.g. Jolly et al. 2000 and Heffener 2004). The second cause has to do with the fact that women perform the majority of the household production, while men specialize in market production (see, e.g., Boye, 2008; Booth and Ours, 2009; Evertsson and Nermo, 2007; Tichenor, 1999). This unequal gender division of household and market work emerges when couples have their first child (Van der Lippe and Siegers, 1994; Sanchez and Thomson 1997; Gauthier and Furstenberg, 2002; Gjerdingen and Center 2005; Baxter et al., 2008). It is possible that women's dual commitments may be more demanding when having a child at older ages, e.g. due to less good health, which then could cause an early retirement.

Let $S_{0i} \equiv 1985 - T_{2i}$ and $S_{1i} \equiv 2010 - T_{2i}$ be the left and right-censoring points for Y_i^*, and let Q_{0i} and Q_{1i} be the corresponding non-censoring indicators:

$$Q_{0i} \equiv 1[S_{0i} \leq Y_i^*] \quad \text{and} \quad Q_{1i} \equiv 1[Y_i^* \leq S_{1i}]; \tag{1}$$

Y_i^* with $Q_{0i}Q_{1i} = 1$ is fully observed.

To summarize, the observed versions of (W_i^*, Y_i^*) are (W_i, Y_i):

$$\begin{aligned} W_i &= DW_i^* + (1-D_i)(45 - T_{2i} + T_{bi}) = \min(W_i^*,\ 45 - T_{2i} + T_{bi}), \\ Y_i &\equiv (1-Q_{0i})S_{0i} + Q_{0i}Q_{1i}Y_i^* + (1-Q_{1i})S_{1i} = \max\{S_{0i}, \min(Y_i^*, S_{1i})\}. \end{aligned}$$

For each woman, what is observed is

$$D_i, W_i,\ S_{0i}, S_{1i}, Q_{0i}, Q_{1i}, Y_i \text{ and } T_{si}\ (= T_{bi} + 45) \text{ along with covariates.}$$

3.2 Identification

3.2.1 Nonparametric Causal Effect Identification without Censoring Problems

Let Y^w be the 'potential duration until retirement from T_2' with the third-child waiting duration $W^* = w$ from T_2; the 'potential duration until retirement from T_2' without third child is Y^0. Assume that the third child dummy $D = D(Z, X, \varepsilon)$ is determined by instruments Z, covariates X observed at T_2, and an error term ε. Since $D = 1[W^* < 45 - T_2 + T_b]$ (i.e., the third-child waiting duration from T_2 does not go over the child-bearing age limit 45) with T_2 and T_b being part of X, we have $W^* = W^*(Z, X, \varepsilon)$.

Let both Y^w and Y^0 be determined by X and an error term U^*:

$$Y^w = Y^w(X, U^*) \quad \text{and} \quad Y^0 = Y^0(X, U^*).$$

The instruments Z does not appear here to satisfy the inclusion and exclusion restrictions (i.e., Z affects W^* and thus D, but not the potential responses). Assume

$$U^* \perp\!\!\!\perp Z|(X, \varepsilon) \implies (Y^w, Y^0) \perp\!\!\!\perp Z|(X, \varepsilon) \implies (Y^w, Y^0) \perp\!\!\!\perp (W^*, D)|(X, \varepsilon),$$

where $\perp\!\!\!\perp$ denotes statistical independence (Dawid 1979). This is a 'selection-on-unobservable' assumption because ε is not observed; ε is to be identified eventually though using the W^* equation.

With W^* as a cardinal treatment and $D = 0$ as the "control treatment", define the individual treatment effect as

$$Y^w - Y^0$$

which is the potential retirement duration difference of a woman having a third child after the waiting time $W^* = w$ versus having no third child. As this is never identified, we strive to identify instead

$$E(Y^w - Y^0).$$

To identify this mean effect from the realized duration Y^*, suppose Y^* and W^* are discrete. Observe

$$
\begin{aligned}
& E(Y^*|X,\varepsilon,\ W^*=w)-E(Y^*|X,\varepsilon,\ D=0) \\
= \ & E(Y^w|X,\varepsilon,\ W^*=w)-E(Y^0|X,\varepsilon,\ D=0) \\
= \ & E(Y^w-Y^0|X,\varepsilon) \quad \{\text{due to } (Y^w,Y^0) \perp\!\!\!\perp (W^*,D)|(X,\varepsilon)\}:
\end{aligned}
$$

the treatment effect conditional on (X,ε) is identified. With $F_{X,\varepsilon}$ as the distribution of (X,ε), we have then

$$
\begin{aligned}
& \int E(Y^*|X=x,\varepsilon=e,W^*=w)-E(Y^*|X=x,\varepsilon=e,\ D=0)dF_{X,\varepsilon}(x,e) \\
= & \int E(Y^w-Y^0|X=x,\varepsilon=e)dF_{X,\varepsilon}(x,e)=E(Y^w-Y^0).
\end{aligned}
$$

3.2.2 Parametric Causal Effect Identification

It would be ideal to identify causal parameters of interest nonparametrically, as just explained. But given the complexity of the issues we face, many covariates to control for, and most importantly, the censoring problems on both Y^* and W^*, we adopt a parametric approach in this paper, which is laid out here.

Let the potential logged durations obey linear models as in

$$
\begin{aligned}
\ln Y^w &= \beta_d+\beta_w w+\beta_{wx}' wX+\beta_x' X+U^* \quad \text{for } w<45-T_2+T_b; && (2) \\
\ln Y^0 &= \beta_x' X+U^* \quad \text{for } w \geq 45-T_2+T_b && (3)
\end{aligned}
$$

where β' are parameters. We use logged durations, because error terms in logged duration tend to be symmetric and homoskedastic, which is more amenable to our parametric approach below.

From (2) and (3), we have

$$
\begin{aligned}
& \ln Y^w-\ln Y^0=\beta_d+\beta_w w+\beta_{wx}' wX \\
\Longrightarrow \quad & E(\ln Y^w-\ln Y^0)=\beta_d+\beta_w w+\beta_{wx}' wE(X) && (4)
\end{aligned}
$$

which is the mean proportional effect, because

$$E\{\ln(Y^w/Y^0)\} = E[\ln\{1+(Y^w-Y^0)/Y^0\}] \simeq E\{(Y^w-Y^0)/Y^0\}.$$

The above mean proportional effect is the effect on the population. For our treatment, it would be more interesting to look at 'the effect on the treated'

$$\begin{aligned} & E(\ln Y^w - \ln Y^0 \ |w < 45 - T_2 + T_b) = E(\beta_d + \beta_w w + \beta_{wx}' wX \ |w < 45 - T_2 + T_b) \\ & = \beta_d + \beta_w w + \beta_{wx}' w \cdot E(X|w < 45 - T_2 + T_b). \end{aligned} \tag{5}$$

Varying w, we can identify $E(Y^w - Y^0|w < 45 - T_2 + T_b)$ over a range of chosen w values. This is the estimand of interest in this paper.[10]

3.3 Estimation

Turning to estimation with observed data (recall $D \equiv 1[W^* < 45 - T_2 + T_b]$), the realized logged work duration is

$$\begin{aligned} \ln Y^* &= (1-D)(\beta_x' X + U^*) + D(\beta_d + \beta_w W + \beta_{wx}' WX + \beta_x' X + U^*) && (6) \\ &= \beta_d D + \beta_w DW + \beta_{wx}' DWX + \beta_x' X + U^*. && (7) \end{aligned}$$

It is important to see that, going from (2) and (3) to (7), *whereas w in (2) is a fixed constant, W in (7) is a random variable: when a random variable is introduced, we should specify how it is generated in relation to the other random variables in the model*, which we do as follows.

For the waiting duration, we assume

$$\ln W^* = \alpha_x' X + \alpha_z' Z + \varepsilon, \qquad \varepsilon \perp\!\!\!\perp (X, Z) \tag{8}$$

[10] Note we can in fact identify each individual effect in (4), with which even the 'α-quantile proportional effect on the treated' $Q_\alpha(\ln Y^w - \ln Y^0|w < 45 - T_2 + T_b)$ can be found:

$$Q_\alpha(\ln Y^w - \ln Y^0|w < 45 - T_2 + T_b) = \beta_d + \beta_w w + Q_\alpha(\beta_{wx}' wX|w < 45 - T_2 + T_b).$$

Bear in mind $Q_\alpha(\beta_{wx}' wX|w < 45 - T_2 + T_b) \neq \beta_{wx}' w Q_\alpha(X|w < 45 - T_2 + T_b)$, because quantile functions are not additive.

where α's are parameters. For the outcome equation, we assume

$$
\begin{aligned}
& U^* = \lambda_u(\varepsilon, X; \rho_{ux}) + U, \quad U \perp\!\!\!\perp (X, Z, \varepsilon) \\
\Longrightarrow \quad & \ln Y^* = \beta_d D + \beta_w DW + \beta_{wx}' DWX + \beta_x' X + \lambda_u(\varepsilon, X; \rho_{ux}) + U
\end{aligned}
$$

where ρ's are parameters and U^* consists of two parts: one part related to (ε, X) and the other part U independent of the first part. $\lambda_u(\varepsilon, X; \rho_{ux})$ is a polynomial functions such as

$$
\lambda_u(\varepsilon, X; \rho_{ux}) = \rho_{\varepsilon 1}\varepsilon + \rho_{\varepsilon 2}\varepsilon^2 + \rho_{\varepsilon 3}\varepsilon^3 + \rho_{\varepsilon x}' X\varepsilon, \quad \rho_{ux} \equiv (\rho_{\varepsilon 1}, \rho_{\varepsilon 2}, \rho_{\varepsilon 3}, \rho_{\varepsilon x}')' \tag{9}
$$

which is just an example, as the exact form to be used in estimation may differ from this.

Since ε appears in the W^* and Y^* equations, (D, W) can be endogenous through ε if ε is unaccounted for. We remove this channel of endogeneity by estimating ε in the W^* equation and then using its estimator $\hat{\varepsilon}$ and its functions as regressors in the Y^* equation; this is a control function (CF) approach. There are several ways to deal with an endogeneity problem in censored models (see Lee 2012, and references therein), but CF approach seems to be the most recommended (see Terza et al. 2008, Kang and Lee 2010, and references therein).

Because the variation of W^* conditional on X and ε stems from the variation of Z due to the inclusion/exclusion restrictions for Z, and because $U \perp\!\!\!\perp (Z, X, \varepsilon)$, we have

$$
U \perp\!\!\!\perp (D, DW)|(X, \varepsilon) \Longleftrightarrow U \perp\!\!\!\perp (D, DW^*)|(X, \varepsilon) \Longrightarrow (Y^w, Y^0) \perp\!\!\!\perp (D, DW^*)|(X, \varepsilon).
$$

Compare this to $U^* \perp\!\!\!\perp Z|(X, \varepsilon) \Longrightarrow (Y^w, Y^0) \perp\!\!\!\perp (W^*, D)|(X, \varepsilon)$ that appeared for nonparametric identification; with X and ε given, the variation of U^* comes U. This is the sense in which the treatment variables (D, DW^*) are independent of (i.e., as good as randomized for) the potential responses (Y^w, Y^0) given the covariates X and CF ε.

If we apply a semiparametric estimator to the $\ln W^*$ equation (8), then ε is estimable only for the $D = 1$ subsample and we can then estimate the $\ln Y^*$ equation accordingly only for this subsample. Consequently D cannot be used as a regressor in the $\ln Y^*$ equation, and the slope of D is not identified. That is, if we are to use a semiparametric estimator for $\ln W$ and to apply the semiparametric procedures suggested in Lee et al. (1996), Chernozhukov and Hong (2002) and Chernozhukov et al. (2015), β_d in the above is not identified. This is unfortunate, because β_d is the parameter of "first-order importance" compared with the rest of the parameters. We therefore proceed parametrically and to impose $\varepsilon \sim N(0, \sigma_\varepsilon^2)$ and $U \sim N(0, \sigma_u^2)$. This implies that we estimate

$$\ln W = \min\{\alpha_x' X + \alpha_z' Z + \varepsilon,\ \ln(45 - T_2 + T_b)\}, \varepsilon \sim N(0, \sigma_\varepsilon^2), \tag{10}$$

where $Z = (1[\text{both boys}], 1[\text{both girls}])$. One caveat about Z is that $Z = 1$ may imply a lower cost in raising the children as they can share clothes and books. However, while this might be an issue in some developing countries, we do not think that it is relevant in the Swedish context. This makes it possible to construct ε for the $D = 0$ sample with 'generalized residuals' as follows.

Letting $\hat{A} \equiv \ln(45 - T_2 + T_b) - \widehat{\alpha}_x' X - \widehat{\alpha}_z' Z$, we have:

$$r_1 \equiv E(\varepsilon | \varepsilon > \hat{A}) = \hat{\sigma}_\varepsilon \frac{\phi(-\hat{A}/\hat{\sigma}_\varepsilon)}{\Phi(-\hat{A}/\hat{\sigma}_\varepsilon)}$$

$$r_2 \equiv E(\varepsilon^2 | \varepsilon > \hat{A}) = \hat{\sigma}_\varepsilon^2 \left[1 + \frac{\hat{A}}{\hat{\sigma}_\varepsilon} \frac{\phi(-\hat{A}/\hat{\sigma}_\varepsilon)}{\Phi(-\hat{A}/\hat{\sigma}_\varepsilon)}\right],$$

and

$$r_3 \equiv E\{\varepsilon^3 | \varepsilon > \hat{A}\} = \hat{\sigma}_\varepsilon^3 \frac{\phi(-\hat{A}/\hat{\sigma}_\varepsilon)}{\Phi(-\hat{A}/\hat{\sigma}_\varepsilon)} \left[2 + \left(\frac{\hat{A}}{\hat{\sigma}_\varepsilon}\right)^2\right]$$

where $\hat{\alpha}_x$, $\hat{\alpha}_z$ and $\hat{\sigma}_\varepsilon$ are the MLE for the $\ln W$ equation. The derivation of r_1, r_2 and r_3 is provided in Appendix A. Hence, for $D_i = 1$, we use $\hat{\varepsilon}_i \equiv \ln W_i - \widehat{\alpha}_x' X - \widehat{\alpha}_z' Z$, $\hat{\varepsilon}_i^2$ and $\hat{\varepsilon}_i^3$ (as well as $\hat{\varepsilon}_i X_i$) as CF, and for $D_i = 0$, we use r_{1i}, r_{2i} and r_{3i} (as well as $r_{1i} X_i$) as CF.

Our second-stage estimation is MLE to the double-censored regression:

$$\ln Y = \max[\ln S_0, \min\{\beta_d D + \beta_w DW + \beta_{wx}' DWX + \beta_x' X \lambda_u(\varepsilon, X; \rho_{ux}) + U, \ln S_1\}\ ],$$

where $S_0 = 1985 - T_2, S_1 = (2010 - T_2), U \sim N(0, \sigma_u^2)$.

If we specify $\lambda_u(\widehat{\varepsilon}, X; \rho_{ux})$ as linear in the parameters as in (9), we estimate

$$\beta_d, \beta_w, \beta_{wx}, \beta_x, \rho_{ux}, \sigma_u$$

in the seceond-stage MLE. The impact of the first-stage estimation error is accounted for by bootstrap.

Pay attention to that we assumed only the marginal normality for both ε and U, not the joint normality of (ε, U^*). The two assumptions are markedly different. If (ε, U^*) is jointly normal, then '$U^* = \rho_\varepsilon \varepsilon + \rho_u U$ with $\varepsilon \perp\!\!\!\perp U$' holding always for some constants ρ_ε and ρ_u. This means that we can set $\lambda_u(\varepsilon, X; \rho_{ux}) = \rho_\varepsilon \varepsilon$ to ignore terms like ε^2, but in our empirical analysis, we find $\hat{\varepsilon}^2$ as well as $\hat{\varepsilon}X$ highly significant. Hence we imposed the weaker marginal normality of ε and U, not the stronger joint normality of (ε, U^*).

4 Data and sample choice

We construct the sample using the multi-generational register from Statistics Sweden (*Flergenerationsregistret*, see SCB (2012)), providing a link between children and their parents. We have information about children born in 2007 at the latest and the observation unit in the present study is a mother with at least two children. To start with, we therefore sample mothers to second-born children from the register. To those mothers, we link information on yearly labor income and education from the longitudinal yearly tables LISA/LOUISE from Statistics Sweden. Coverage for education and labor income is universal. Education is measured in theoretical years of education using the official Swedish SUN classification, which roughly follows the international ISCED 97 standard.

We use labor income to estimate the year of retirement. In particular, if a mother has an income below a specific threshold during three consecutive years, she is regarded as retired starting from the first of those three years. We have yearly income data for the years 1985–2010. Information about labor market income is based on the annual reports from the employers to the tax authorities. Thus, this amount includes the total individual pre-tax income from work, and does not include the amount of paid parental leave, tax reductions, or social transfers such as means-tested cash benefits. The income measure does however include a part of the sick-pay: In case of illness, the first day is not replaced. Thereafter, the employer pays sick-pay for the 13 following days, and this amount is included in the income measure.

The threshold determining retirement is a half, so called, *price base amount* for the relevant income year. A price base amount tracks inflation and was SEK 44,500 (about EUR 4,400) in 2015. The calculation of the Swedish price base amount is regulated by law and used for various types of benefits, insurance payments, etc. It is higher than the mean monthly wage (SEK 29,200 or EUR 2,900 for women in 2014). Thus, using half of the price base amount as a threshold is equivalent to using roughly 76 percent of the mean monthly wage.

We choose the sample using the following criteria (with the resulting number of observations within parentheses):

1. We start by sampling mothers who were born between 1923 and 1963 (1,607,906 obs.).
2. The mothers were at most 44 years old when giving second or third birth (1,605,905 obs.).
3. No twins at first or second birth, but possible twins at third birth or higher (1,572,932 obs.).
4. The durations between first and second birth, and between second and third birth, should be longer than 7 months (1,572,745 obs.).
5. Retirement is at or after the end of fertility, i.e, age 45 (1,292,280 obs.).
6. Individuals with missing information on education are removed from the sample (1,261,855 obs.).
7. As a final step, we remove mothers born later than 1947 (804,721 obs.).

Table 1: Number of observations by year of birth

	1923–24	1925–29	1930–34	1935–39	1940–44	1945–47
Number of observations	56,701	142,389	146,188	160,961	179,508	118,974

The number of observations by birth year are presented in Table 1. The relative numbers of observations reflect the corresponding demographics in the Swedish population as a whole. For instance, the higher numbers of mothers from the baby boom-generations during the 1940's are clearly seen in the table.

The reasoning behind (1) is explained in detail in section 3. Conditions (2) and (4) are imposed because we want to exclude retirement for potential health reasons, and (3) gives a precise fertility measure. Further, conditions (1) and (2) together with the fact that we have data on children born in 2007 at the latest implies that our measure of completed fertility (as per 2007) is measured when the youngest cohort (those born in 1963) were aged 44. In theory, this implies that some of the youngest mothers might have given a second or third birth after the age of 44 (i.e., in the year 2008, 2009 or 2010, when we have income data). In such case, those mothers would have been excluded from the sample due to condition (2), if data on child births after 2007 was available. However, as the share of women giving birth after the age of 44 is very small, and as we do not use these younger cohorts in the end this is not a problem for the present study. Finally, (5) precludes the effect (retirement) from preceding the cause (more than two children). This last condition reduces the sample with about 22 percent. Clearly, if we were interested in the decisions of relatively young women of whether to exit the labor force, this restriction could be problematic. However, in the present paper, we are interested in retirement, and defining retirement as the decision to exit the labor market after 45 but not earlier seems natural.

More importantly, a substantial share of the women excluded from the sample due to (5) have probably never entered the labor market. We do not know the exact share since we lack income data before 1985. In other words, we do not have information on whether a person worked or not before 1985. However, we know from official statistics that about

25 percent of the women between 15 and 44 years are not in the labor force.[11] In light of this figure, we are not too worried about the 22 percent drop in sample size as a result of condition (5) above. It is likely that most of the observations dropped due to the sample choice condition had very low labor market attachment, or never took part in the labor force to start with.

The sample choice criteria 1–6 imply that the share of right-censored observations with respect to retirement is about 43 percent. The reason for this is that we include relatively young cohorts, and although there is a theoretical possibility for younger mothers to retire within the data window, few mothers in Sweden retire at an early age. To deal with this, in the final selection stage 7, we remove mothers born later than 1947. This results in the same share of right- and left-censored observations (16 percent). These women born 1947 are hence 23 years old in 1970. The later cohorts will to a larger degree be affected by the changes in institutions with regards to taxation, parental leave and child care described in section two. We therefore in a sensitvity analysis redo the analysis for the sample of 374,932 women born between 1923-1935 . For these cohorts the institutions should be fairly constant for the population under study.

The data are described in Table 2 and we start by presenting the first column, containing means and standard deviations for the whole sample. The average number of years of education is slightly above 10 years, which is approximately one year longer than primary school. Further, the mothers in the sample were on average about 27.5 years old when giving birth the second time. Half of the mothers had two first children of the same sex, of which 26 percent were two boys and 24 percent were two girls. Of the mothers of at least two children that constitute the sample, 43 percent had at least three children, and for those mothers, the spacing between their second and third child was on average 4.4 years.

We now go on to describe the sample by treatment status (columns two and three in Table 2). First, as expected, mothers with shorter education have higher number of

[11] Source: The Swedish Labor Survey, AKU (2005), available on the internet. The year 2005 is the earliest available, and the corresponding number for 2014 was 23 percent. The lower age limit of 15 years is not our choice but the way Statistics Sweden organize their aggregate data presentations.

Table 2: Descriptive statistics

	All	D = 1	D = 0	same = 1	same = 0
Years of education (*educ*)	10.20	9.97	10.39	10.20	10.21
	(3.03)	(3.02)	(3.02)	(3.03)	(3.03)
Age at 2nd birth (*age2nd*)	27.52	25.58	29.00	27.53	27.51
	(4.58)	(3.86)	(4.54)	(4.58)	(4.58)
1[Same sex] (*same*)	0.50	0.53	0.48		
	(0.50)	(0.50)	(0.50)		
1[Two boys] (*boys*)	0.26	0.28	0.25	0.53	
	(0.44)	(0.45)	(0.43)	(0.50)	
1[Two girls] (*girls*)	0.24	0.25	0.23	0.47	
	(0.42)	(0.43)	(0.42)	(0.50)	
1[> 2 children] (D)	0.43			0.46	0.41
	(0.50)			(0.50)	(0.49)
Spacing 3rd-2nd | > 2 (DW)*	4.38	4.38		4.32	4.46
	(3.06)	(3.06)		(2.98)	(3.14)
1[Y left censored] (C_0)	0.16	0.19	0.14	0.16	0.16
	(0.37)	(0.39)	(0.34)	(0.37)	(0.37)
1[Y right censored] (C_1)	0.16	0.13	0.18	0.16	0.16
	(0.37)	(0.34)	(0.39)	(0.37)	(0.37)
Number of observations	804,721	348,681	456,040	401,990	402,731

Notes: Standard deviations within parentheses and variable names used throughout the paper *in italics*. Education is measured in theoretical years of education using the official Swedish SUN classification, which roughly follows the international ISCED 97 standard.
*Means and standard deviations for DW are calculated only for individuals where $D = 1$. This implies that the samples used to calculate DW are subsamples of the corresponding groups, except for column two where $D = 1$ already.

children: mothers with two children have 10.4 years of education on average and the corresponding number for mothers with more than two children is 10.0. Also, as expected, mothers with two children give second birth on average 3.42 ($= 29 - 25.58$) years later compared to mothers with more than two children.

Comparing the share with at least three children among mothers to two children of the same sex (53 percent) with the corresponding share among mothers of two mixed-sex children (48 percent) gives an approximation of the relevance of the 'same-sex' instrument: The likelihood of having more than two children is about 5 percentage points higher among mothers whose two children have the same sex, compared to mothers of children of mixed sex. This difference is of the same magnitude (about one percentage point lower) as the corresponding number reported in Angrist and Evans (1998).

As in previous studies, having two boys is associated with a greater difference in the likelihood of having more than three children compared to having two girls (a difference of three percentage points for those with two boys and two percentage points for those with two girls). Qualitatively, this difference is in line with previous studies using the same-sex instrument. Finally, there is some difference in the share of censored observations among those with two children and those with more than two children: A larger share of the mothers with more than two children have left-censored durations on the labor market, and a larger share of the mothers with two children have right-censored durations.

Finally, we present descriptives by the value of the same-sex instrument in the two final columns of Table 2. Here, it is important that years of education and age at second birth do not differ between mothers with same sex- or mixed-sex children, and the table reveals no differences.

5 Results

In the estimation we control for education (*educ*), age at second birth (*age2nd*), and birth year of the mother (*birth year*). With regard to the empirical specification of $h^p(D,W,X;\beta_{dwx})$ we have chosen to categorize W according to its quartiles and interacted W with *age2nd* and *educ*. Hence,

$$\begin{aligned} h^p(D_i,W_i,X_i;\beta_{dwx}) = {} & \beta_d D_i + \beta_{w2} 1_{q_2} + \beta_{w3} 1_{q_3} + \beta_{w4} 1_{q4} + \\ & \beta_{as} as_i D_i + \beta_{as2} 1_{q_2} as_i + \beta_{as3} 1_{q_3} as_i + \beta_{as4} 1_{q4} as_i + \\ & \beta_{ed} ed_i D_i + \beta_{ed2} 1_{q_2} ed_i + \beta_{ed3} 1_{q_3} ed_i + \beta_{ed4} 1_{q_4} ed_i, \end{aligned}$$

where $1_{q_2} = [Q_1 < W_i \le Q_2], 1_{q_3} = [Q_2 < W_i \le Q_3], 1_{q_4} = [Q_3 < W_i]$, $(Q_1,Q_2,Q_3) = (2,3.5,6)$ are the quartiles of W measured in years when $D = 1$, $ed_i = (educ_i\text{-}\overline{educ})$ and $as_i = (age2nd - \overline{age2nd})$, where denotes the sample mean, where $\overline{age2nd} = 27.5$ and $\overline{educ} = 10.2$.

We have throughout interacted the 'fitted residuals' $D\widehat{\varepsilon}+(1-D)r_1$ with *educ*, *age2nd* and *birth year* and besides the first moment, we have included second and third moments (i.e., $D\widehat{\varepsilon}^2+(1-D)r_2$ and $D\widehat{\varepsilon}^3+(1-D)r_3$ as well).

The estimation results are presented in Table 3, where we have included the first stage of the estimation in column (1) along with the final stage in columns (2) and (3). Throughout, we use the separate instruments $boys = 1[\text{ first two children are boys}]$ and $girls = 1[\text{first two children are girls}]$. From the first stage in the first column of Table 3, we see that the instruments are relevant for the spacing between the second and third child since the coefficients for *boys* and *girls* are statistically significant when we condition on the rest of the covariates. We present a Wald-test of joint significance at the bottom of the table. The χ^2_2-statistic is 2389.7 and the corresponding F-value is 1194.9. Having two firstborn boys or girls is associated with approximately 20 percent shorter spacing between the second and third child. It is of interest to note that conditional on age at second birth, and birth cohort, women with higher education are more likely to have a third child earlier.[12] This result is not unexpected given that highly educated mothers give birth later than less educated. Hence, once we control for the age at second birth higher educated mothers are more likely to have a third child. This may reflect budget restrictions affecting the choice of having a third child but also differences in health etc. This result is interesting in itself given the previous, often unconditional, descriptive statistics of fertility and education indicating a negative sorting.

The results from the estimation where we neglect potential endogeneity are displayed in column (2) of Table 3, while the main results are displayed in column (3). The parameter estimates from column (2) suggest a reduction of time on the labor market if having a third child (by -3.2 percent) for the average (i.e. $educ = 10.2$ years and $age2nd = 27.5$ years) mother with the shortest birth spacing (i.e. $W \leq 2$ years) between second and third child. This negative estimate is monotonously decreasing in absolute value towards zero for longer waiting times. For the women with the longest waiting time (more than 6 years) there is no reduction in labor market work duration. Turning to the main results displayed

[12]Conditional on age at second birth and birth cohort, women with higher education are also more likely to have a third child. These results are available upon request.

in column (3) one can see that the parameters for the control function (CF) terms as well as their interaction are individually statistically significant. The CF-parameters and their interaction are also jointly statistically significant (see the Wald-test at the bottom of the table). This highlights the fact that having a third child depends on future labor market prospects.

The estimates from the CF-specification give a very different picture than from what was displayed in column (2). We find an overall reduction in the labor market duration with about 5.2 percent for an average mother with a child spacing less than or equal to two years. As the average duration on the labor market, Y, is about 34 years[13], the point estimate for D implies that having a third child within two years after the second child shortens the duration on the labor market with about 1.8 year. Now, child spacing longer than one year reduces labor market work duration and the magnitude of the negative effect increases monotonically with child spacing. The effect for the longest child spacing is -7.9 percent compared to -5.2 percent for the shortest child spacing. These child spacing effect differs from what was found in Karimi (2014). She found that a one year delay of two child mothers increased the labor income up to fifteen years after the parenthood. It is difficult to have an opinion about the differences as we study two different populations (one child mothers against two child mothers), differences in outcomes (income vs duration) and take use of different instruments (miscarriages vs same sex). However, it should be clear that Karimi estimates the marginal effects of a one year delay of parenthood irrespective of the birth spacing length while we estimate effects for a given birth space against less than two years birth space. The miscarriage shifts the distribution of child spacing after two years and therefore the marginal effect found in Karimi (2014) does not capture the effects of less than two years birth spacing against three years of birth spacing.

From the interaction terms of D and the quantile-categorized DW, we can see that both the effect of having a third child and that of child spacing depend on the age of the mother at second child birth and to some extent also on mothers' education level. In order to facilitate the interpretation of the interaction terms, we present a counterfactual analysis in Tables 4–7 using the results from Table 3. There, we have calculated the effect of

[13] Recall that Y is censored, but the share of RC and LC observations is equal.

Table 3: Estimation results

	Dependent variable:		
	ln(w) (right-censored)	y (double-censored)	
	CF	MLE	MLE with CF
D		−0.032***	−0.052***
		(0.001)	(0.009)
$1[Q_1 < W \leq Q_2]$		0.012***	−0.017***
		(0.001)	(0.002)
$1[Q_2 < W \leq Q_3]$		0.025***	−0.021***
		(0.001)	(0.003)
$1[Q_3 < W]$		0.029***	−0.027***
		(0.002)	(0.004)
$educ$	−0.041***	0.020***	0.026***
	(0.001)	(0.0001)	(0.0003)
$birth\ year$	0.043***	0.004***	0.004***
	(0.0003)	(0.00005)	(0.0001)
$age2nd$	0.151***	−0.032***	−0.041***
	(0.001)	(0.0001)	(0.0004)
$boys$	−0.225***		
	(0.005)		
$girls$	−0.202***		
	(0.005)		
$D\widehat{\varepsilon} + (1-D)r_1$			1.112***
			(0.107)
$D\widehat{\varepsilon}^2 + (1-D)r_2$			−0.015***
			(0.001)
$D\widehat{\varepsilon}^3 + (1-D)r_3$			−0.001***
			(0.0002)
$educ \times (D\widehat{\varepsilon} + (1-D)r_1)$			−0.006***
			(0.0002)
$birth\ year \times (D\widehat{\varepsilon} + (1-D)r_1)$			−0.001***
			(0.0001)
$age2nd \times (D\widehat{\varepsilon} + (1-D)r_1)$			0.005***
			(0.0002)
$D \times age2nd$		0.004***	0.032***
		(0.0002)	(0.001)
$1[Q_1 < W \leq Q_2] \times age2nd$		0.0005	−0.005***
		(0.0003)	(0.0004)
$1[Q_2 < W \leq Q_3] \times age2nd$		0.002***	−0.009***
		(0.0003)	(0.0004)
$1[Q_3 < W] \times age2nd$		0.002***	−0.016***
		(0.0004)	(0.001)
$D \times educ$		0.003***	−0.019***
		(0.0003)	(0.001)
$1[Q_1 < W \leq Q_2] \times educ$		−0.0002	0.004***
		(0.0004)	(0.0004)
$1[Q_2 < W \leq Q_3] \times educ$		−0.001*	0.007***
		(0.0004)	(0.0005)
$1[Q_3 < W] \times educ$		−0.001***	0.011***
		(0.0005)	(0.001)
$Intercept$	−79.449***	−4.587***	−3.832***
	(0.625)	(0.091)	(0.173)
	Wald-tests for joint parameter significance		
χ^2_2 for $(boys, girls)$	2389.7		
χ^2_5 for CF and interactions			4191.6
Log Likelihood	−938,352.6	−239,804.5	−236,555.9
No. of observations	804,721	804,721	804,721

Notes: *p<0.1; **p<0.05; ***p<0.01. Bootstrapped standard errors in parentheses (1,000 replications).
† Q_1, Q_2, Q_3 are the quartiles of W for the part of the distribution where $D = 1$. The parameters for the three categorical variables are estimated relative to the reference $1[0 < W \leq Q_1]$.

having a third child on labor market duration for different child spacing groups, ages at second child birth and years of education. In the estimation we interacted age at second birth and years of education linearly with *D* and the categorical variables for child spacing. This should be borne in mind when looking at the results from the counterfactual analyses, in the sense that the exact magnitude of effect estimates at each age and year of education cell should be interpreted with a grain of salt. However, note that the number of individuals in each cell is large so we are not extrapolating outside of our data.

From tables 4–7, it becomes clear that the effect is monotonically decreasing with age at second child and increasing with years of education. The results also reveal that the range of the effect variation over age at second birth is monotonically decreasing with child spacing. For instance, for low educated mothers the effect estimates of child spacing across the age distribution (i.e. 22 years compared with 34 years) in the lowest part of the child spacing distribution is in the range -0.208 to 0.181, or a difference of about 39 percentage points (see the first column in Table 4). The corresponding effect estimates in the upper-most part of the spacing distribution is in the range -0.160 to 0.036, or a difference of about 16.4 percentage points (see the first column in Table 7).

With respect to heterogeneous effects over education, for the lowest part of the spacing distribution, the effect on labor market duration is around 13 percentage points larger in magnitude for those with 16 years of education compared with those with 9 years of education among the youngest mothers (-0.338 and -0.208, respectively, as seen in the first row of Table 4). The corresponding effects higher up in the child spacing distribution show a 10.2 (between the first and second quartile), 8.2 (between the second and third quartile), and 5.1 (above the third quartile) percentage points higher duration effect for the highest educated compared to the lowest educated mothers (see row one in Tables 5–7). For the oldest mothers, the corresponding reductions for the highest educated in comparison with the lowest are 13.1, 10.2, 9.1, and 5.1 percentage points over the four different parts of the child spacing distribution, respectively.

Table 4: The effect of having a third child on labor market duration for mothers with $0 < DW \leq Q_1$

Age at second birth	Years of education: 9	11	12	14	16
22	-0.208***	-0.245***	-0.264***	-0.301***	-0.338***
	(0.010)	(0.010)	(0.010)	(0.009)	(0.009)
	689	*2735*	*205*	*305*	*170*
23	-0.175***	-0.213***	-0.231***	-0.269***	-0.306***
	(0.009)	(0.010)	(0.010)	(0.010)	(0.009)
	691	*2893*	*242*	*336*	*252*
24	-0.143***	-0.180***	-0.199***	-0.236***	-0.274***
	(0.009)	(0.009)	(0.009)	(0.009)	(0.009)
	698	*2868*	*287*	*375*	*408*
25	-0.110***	-0.148***	-0.167***	-0.204***	-0.241***
	(0.009)	(0.009)	(0.009)	(0.009)	(0.009)
	692	*2599*	*306*	*545*	*581*
26	-0.078***	-0.116***	-0.134***	-0.172***	-0.209***
	(0.009)	(0.009)	(0.009)	(0.009)	(0.008)
	570	*2306*	*315*	*624*	*786*
27	-0.046***	-0.083***	-0.102***	-0.139***	-0.177***
	(0.008)	(0.008)	(0.008)	(0.008)	(0.008)
	566	*2010*	*340*	*641*	*963*
28	-0.014	-0.051***	-0.070***	-0.107***	-0.144***
	(0.008)	(0.008)	(0.008)	(0.010)	(0.009)
	437	*1647*	*295*	*652*	*1035*
29	0.019**	-0.019**	-0.037***	-0.075***	-0.112***
	(0.009)	(0.009)	(0.009)	(0.009)	(0.010)
	366	*1372*	*272*	*559*	*917*
30	0.051***	0.014	-0.005	-0.042***	-0.080***
	(0.010)	(0.009)	(0.009)	(0.009)	(0.009)
	318	*1089*	*204*	*433*	*786*
31	0.084***	0.046***	0.027***	-0.010	-0.047***
	(0.009)	(0.009)	(0.009)	(0.009)	(0.009)
	258	*834*	*168*	*369*	*632*
32	0.116***	0.078***	0.060***	0.022**	-0.015
	(0.009)	(0.009)	(0.009)	(0.008)	(0.009)
	172	*614*	*115*	*268*	*492*
33	0.148***	0.111***	0.092***	0.055***	0.017*
	(0.009)	(0.009)	(0.008)	(0.008)	(0.008)
	134	*483*	*98*	*216*	*378*
34	0.181***	0.143***	0.124***	0.087***	0.050***
	(0.008)	(0.008)	(0.008)	(0.008)	(0.008)
	112	*411*	*63*	*131*	*283*

Notes: *p<0.1; **p<0.05; ***p<0.01. Bootstraped standard errors in parentheses and number of observations in each cell in italics.

Table 5: The effect of having a third child on labor market duration for mothers with $Q_1 < DW \leq Q_2$

Age at second birth		Years of education				
		9	11	12	14	16
	22	-0.201***	-0.230***	-0.244***	-0.274***	-0.303***
		(0.010)	(0.010)	(0.010)	(0.009)	(0.009)
		560	*2350*	*166*	*263*	*177*
	23	-0.174***	-0.203***	-0.217***	-0.246***	-0.276***
		(0.009)	(0.009)	(0.010)	(0.010)	(0.009)
		673	*2561*	*233*	*359*	*266*
	24	-0.147***	-0.176***	-0.190***	-0.219***	-0.249***
		(0.009)	(0.009)	(0.009)	(0.009)	(0.009)
		695	*2703*	*291*	*431*	*380*
	25	-0.120***	-0.149***	-0.163***	-0.192***	-0.222***
		(0.009)	(0.009)	(0.009)	(0.009)	(0.008)
		686	*2634*	*342*	*556*	*659*
	26	-0.093***	-0.122***	-0.136***	-0.165***	-0.194***
		(0.008)	(0.008)	(0.009)	(0.009)	(0.008)
		681	*2474*	*379*	*660*	*991*
	27	-0.066***	-0.095***	-0.109***	-0.138***	-0.168***
		(0.008)	(0.008)	(0.008)	(0.008)	(0.008)
		511	*2167*	*400*	*821*	*1187*
	28	-0.038***	-0.068***	-0.082***	-0.111***	-0.140***
		(0.008)	(0.008)	(0.008)	(0.010)	(0.010)
		462	*1769*	*341*	*744*	*1278*
	29	-0.012	-0.041***	-0.055***	-0.084***	-0.113***
		(0.010)	(0.009)	(0.009)	(0.009)	(0.010)
		373	*1442*	*285*	*652*	*1186*
	30	0.016*	-0.014	-0.028***	-0.057***	-0.086***
		(0.010)	(0.009)	(0.009)	(0.009)	(0.009)
		287	*1154*	*221*	*546*	*980*
	31	0.043***	0.013	-0.001	-0.030***	-0.059***
		(0.009)	(0.009)	(0.009)	(0.009)	(0.009)
		189	*835*	*177*	*410*	*729*
	32	0.070***	0.040***	0.026***	-0.003	-0.032***
		(0.009)	(0.009)	(0.009)	(0.008)	(0.009)
		167	*585*	*129*	*323*	*608*
	33	0.097***	0.068***	0.053***	0.024**	-0.005
		(0.009)	(0.009)	(0.008)	(0.008)	(0.008)
		113	*493*	*97*	*242*	*366*
	34	0.124***	0.094***	0.080***	0.051***	0.022**
		(0.008)	(0.008)	(0.008)	(0.008)	(0.008)
		89	*302*	*65*	*135*	*310*

Notes: *p<0.1; **p<0.05; ***p<0.01. Bootstraped standard errors in parentheses and number of observations in each cell in italics.

Table 6: The effect of having a third child on labor market duration for mothers with $Q_2 < DW \leq Q_3$

Age at second birth	Years of education: 9	11	12	14	16
22	-0.186***	-0.210***	-0.221***	-0.244***	-0.268***
	(0.010)	(0.010)	(0.010)	(0.010)	(0.010)
	607	*2511*	*203*	*289*	*203*
23	-0.163***	-0.186***	-0.198***	-0.221***	-0.244***
	(0.010)	(0.010)	(0.010)	(0.010)	(0.009)
	774	*3044*	*252*	*348*	*240*
24	-0.140***	-0.163***	-0.175***	-0.198***	-0.221***
	(0.009)	(0.009)	(0.009)	(0.009)	(0.009)
	750	*3118*	*318*	*465*	*363*
25	-0.117***	-0.140***	-0.152***	-0.175***	-0.198***
	(0.009)	(0.009)	(0.009)	(0.009)	(0.009)
	765	*3145*	*385*	*570*	*522*
26	-0.094***	-0.117***	-0.128***	-0.152***	-0.175***
	(0.009)	(0.009)	(0.009)	(0.009)	(0.008)
	742	*3005*	*401*	*672*	*897*
27	-0.070***	-0.094***	-0.105***	-0.128***	-0.152***
	(0.008)	(0.008)	(0.008)	(0.008)	(0.008)
	681	*2575*	*410*	*783*	*1133*
28	-0.047***	-0.070***	-0.082***	-0.105***	-0.128***
	(0.008)	(0.008)	(0.008)	(0.010)	(0.010)
	525	*2135*	*404*	*810*	*1235*
29	-0.024***	-0.047***	-0.059***	-0.082***	-0.105***
	(0.010)	(0.010)	(0.010)	(0.010)	(0.010)
	481	*1686*	*312*	*694*	*1225*
30	-0.001	-0.024***	-0.036***	-0.059***	-0.082***
	(0.010)	(0.010)	(0.010)	(0.009)	(0.009)
	348	*1300*	*225*	*515*	*1033*
31	0.022***	-0.001	-0.012	-0.036***	-0.059***
	(0.009)	(0.009)	(0.009)	(0.009)	(0.009)
	240	*844*	*177*	*396*	*734*
32	0.046***	0.022***	0.011	-0.012	-0.036***
	(0.009)	(0.009)	(0.009)	(0.009)	(0.009)
	159	*599*	*132*	*269*	*523*
33	0.069***	0.046***	0.034***	0.011	-0.012
	(0.009)	(0.009)	(0.008)	(0.008)	(0.008)
	87	*378*	*86*	*159*	*361*
34	0.092***	0.069***	0.057***	0.034***	0.011
	(0.008)	(0.008)	(0.008)	(0.008)	(0.008)
	67	*228*	*50*	*113*	*231*

Notes: *p<0.1; **p<0.05; ***p<0.01. Bootstraped standard errors in parentheses and number of observations in each cell in italics.

Table 7: The effect of having a third child on labor market duration for mothers with $Q_3 < DW$

Age at second birth	Years of education: 9	11	12	14	16
22	-0.160*** (0.011) *611*	-0.174*** (0.010) *2539*	-0.182*** (0.010) *237*	-0.196*** (0.010) *335*	-0.211*** (0.010) *195*
23	-0.144*** (0.010) *689*	-0.158*** (0.010) *2881*	-0.165*** (0.010) *273*	-0.180*** (0.010) *351*	-0.195*** (0.010) *264*
24	-0.127*** (0.010) *687*	-0.142*** (0.009) *2873*	-0.149*** (0.009) *312*	-0.164*** (0.009) *444*	-0.178*** (0.009) *345*
25	-0.111*** (0.010) *717*	-0.126*** (0.009) *2785*	-0.133*** (0.009) *316*	-0.147*** (0.009) *468*	-0.162*** (0.009) *491*
26	-0.095*** (0.009) *629*	-0.109*** (0.009) *2477*	-0.116*** (0.009) *347*	-0.131*** (0.009) *472*	-0.146*** (0.008) *657*
27	-0.078*** (0.008) *521*	-0.093*** (0.008) *2003*	-0.100*** (0.008) *309*	-0.115*** (0.008) *555*	-0.129*** (0.008) *783*
28	-0.062*** (0.008) *400*	-0.077*** (0.008) *1556*	-0.084*** (0.008) *278*	-0.098*** (0.011) *465*	-0.113*** (0.011) *777*
29	-0.046*** (0.010) *349*	-0.060*** (0.010) *1125*	-0.068*** (0.010) *204*	-0.082*** (0.010) *388*	-0.097*** (0.010) *631*
30	-0.029*** (0.010) *245*	-0.044*** (0.010) *781*	-0.051*** (0.010) *150*	-0.066*** (0.010) *270*	-0.080*** (0.010) *502*
31	-0.013* (0.010) *148*	-0.028*** (0.010) *490*	-0.035*** (0.010) *104*	-0.050*** (0.010) *191*	-0.064*** (0.009) *378*
32	0.003 (0.009) *80*	-0.011 (0.009) *278*	-0.019** (0.009) *46*	-0.033*** (0.009) *114*	-0.048*** (0.009) *208*
33	0.020** (0.009) *54*	0.005 (0.009) *176*	-0.002 (0.008) *32*	-0.017** (0.008) *71*	-0.032*** (0.008) *135*
34	0.036*** (0.008) *29*	0.021*** (0.008) *80*	0.014* (0.008) *15*	-0.001 (0.008) *34*	-0.015* (0.008) *67*

Notes: *p<0.1; **p<0.05; ***p<0.01. Bootstraped standard errors in parentheses and number of observations in each cell in italics.

5.1 Results for the cohorts 1923-1935

To obtain some understanding of the importance of the institutional changes in the 1970's for the results we redo the analysis for the cohorts born between 1923-1935. Using these cohorts, we ensure that the absolute majority of mothers were unaffected by the new policies, since the youngest mothers were aged 35 in 1970. The results from the estimation are presented in Table 8.

The results are qualitatively very similar to the previous ones and we therefore focus the discussion on the main results displayed in column (3). The overall reduction in the labor market duration for an average mother with a child spacing less than or equal to two years is lower than in the main analysis: a recuction by 3.2 percent compared to 5.2 percent. As the average duration on the labor market is about 33.65 years for these cohorts, this implies that having a third child within two years after the second child shortens the duration on the labor market with about 1.1 year. Child spacing longer than two years leads to a further reduction in labor market work duration. The magnitude of the negative effect increases monotonically with child spacing. The effect for the longest child spacing is -4.4 percent compared to -3.2 percent for the shortest. From the interaction terms of D and the quantile-categorized DW, we can once again see that both the effect of having a third child and that of child spacing depend on the age of the mother at second child birth and to some extent also on mothers' education level. In order to facilitate the interpretation of the interaction terms, we once again present the counterfactual analyses in Tables 9–12.

From these tables we can see that for low educated mothers the effect estimates of child spacing across the age distribution (i.e. 22 years compared with 34 years) in the lowest part of the child spacing distribution is in the range -0.201 to 0.124, or a difference of about 32 percentage points (see the first column in Table 9) which can be compared to a 39 percentage points difference in the main analysis. The corresponding effect estimates in the upper-most part of the spacing distribution is in the range -0.130 to 0.020, or a difference of about 11 percentage points (see the first column in Table 12). The corresponding difference in the main analysis was 16.4 percentage points.

Table 8: Estimation results for the restricted sample of mothers born 1923–1935

	Dependent variable:		
	ln(w) (right-censored)	y (double-censored)	
	CF	MLE	MLE with CF
D		−0.020***	−0.032***
		(0.001)	(0.009)
$1[Q_1 < W \leq Q_2]$		0.005***	−0.009***
		(0.001)	(0.002)
$1[Q_2 < W \leq Q_3]$		0.011***	−0.010***
		(0.001)	(0.003)
$1[Q_3 < W]$		0.017***	−0.011***
		(0.001)	(0.004)
educ	−0.033***	0.016***	0.018***
	(0.001)	(0.0001)	(0.0003)
birth year	0.018***	−0.005***	−0.005***
	(0.001)	(0.0001)	(0.0001)
age2nd	0.139***	−0.032***	−0.038***
	(0.001)	(0.0001)	(0.0004)
boys	−0.190***		
	(0.007)		
girls	−0.174***		
	(0.007)		
$D\widehat{\varepsilon} + (1-D)r_1$			−0.690***
			(0.124)
$D\widehat{\varepsilon}^2 + (1-D)r_2$			−0.008***
			(0.001)
$D\widehat{\varepsilon}^3 + (1-D)r_3$			−0.001***
			(0.0002)
$educ \times (D\widehat{\varepsilon} + (1-D)r_1)$			−0.002***
			(0.0002)
$birth\ year \times (D\widehat{\varepsilon} + (1-D)r_1)$			0.0004***
			(0.0001)
$age2nd \times (D\widehat{\varepsilon} + (1-D)r_1)$			0.004***
			(0.0002)
$D \times age2nd$		0.003***	0.022***
		(0.0002)	(0.0005)
$1[Q_1 < W \leq Q_2] \times age2nd$		0.0004	−0.004***
		(0.0003)	(0.0004)
$1[Q_2 < W \leq Q_3] \times age2nd$		0.001**	−0.007***
		(0.0003)	(0.0004)
$1[Q_3 < W] \times age2nd$		0.001***	−0.011***
		(0.0003)	(0.0005)
$D \times educ$		0.003***	−0.003***
		(0.0003)	(0.001)
$1[Q_1 < W \leq Q_2] \times educ$		−0.001	0.001
		(0.0004)	(0.0004)
$1[Q_2 < W \leq Q_3] \times educ$		−0.001***	0.001**
		(0.0004)	(0.0005)
$1[Q_3 < W] \times educ$		−0.003***	0.001***
		(0.0005)	(0.001)
Intercept	−31.381***	13.206***	13.600***
	(1.528)	(0.156)	(0.170)
Log Likelihood	−456,166.700	−2,087.436	−466.588
No. of observations	374,932	374,932	374,932

Notes: *p<0.1; **p<0.05; ***p<0.01. Bootstrapped standard errors in parentheses (1,000 replications).
† Q_1, Q_2, Q_3 are the quartiles of W for the part of the distribution where $D = 1$. The parameters for the three categorical variables are estimated relative to the reference $1[0 < W \leq Q_1]$.

With respect to heterogeneous effects over education, for the lowest part of the spacing distribution, the effect on labor market duration is around 5 percentage points larger in magnitude for those with 16 years of education compared with those with 9 years of education among the youngest mothers (-0.211 and -0.160, respectively, as seen in the first row of Table 9). The corresponding effects higher up in the child spacing distribution show a 1.8 (between the first and second quartile), 1.4 (between the second and third quartile), and 1.1 (above the third quartile) percentage points higher duration effect for the highest educated compared to the lowest educated mothers (see row one in Tables 10–12). The corresponding differences in effects in the main analysis were 13, 10.2, 8.2, and 5.1 percentage points.

All in all, the effects are smaller in magnitude for the older cohorts. These smaller effects are especially pertaining to the heterogeneous effects with respect to schooling. It is difficult to explain these differences in magnitudes. If anything, given that the reforms implemented in 1970's encouraged the labor supply of women, we would have expected the opposite, that is a larger negative effect of a third child on labor for the early cohorts. However, the effects on third child could be different than the effects on parenthood per se. A possible interpretation is that mothers with three children who worked in the early cohorts are a more selective population of women with a better labor market attachment than later cohorts of working mothers.

5.2 Discussion

The question is how all these effects should be interpreted. Suppose that age at second birth given education is a proxy for the unobserved labor market attachment. One argument supporting this is that it is reasonable to assume that labor success delays child bearing. That is, for two women with the same education and skill when leaving school, it is more likely that the one delaying child birth is more successful and committed to her work than the one advancing. The heterogeneous responses across the age distribution then support the idea that there are small long-term effects on the labor market duration from having a third child for mothers with a strong position and/or attachment to the labor market.

Table 9: The effect of having a third child on labor market duration for mothers with $0 < DW \leq Q_1$ for the restricted sample of mothers born 1923–1935

Age at second birth	Years of education: 9	11	12	14	16
22	-0.160***	-0.165***	-0.168***	-0.173***	-0.178***
	(0.009)	(0.009)	(0.009)	(0.009)	(0.009)
	259	*1111*	*54*	*78*	*36*
23	-0.138***	-0.143***	-0.146***	-0.151***	-0.157***
	(0.009)	(0.009)	(0.009)	(0.009)	(0.008)
	306	*1209*	*59*	*86*	*69*
24	-0.116***	-0.122***	-0.124***	-0.130***	-0.135***
	(0.008)	(0.008)	(0.008)	(0.008)	(0.009)
	320	*1216*	*81*	*81*	*106*
25	-0.095***	-0.100***	-0.103***	-0.108***	-0.113***
	(0.009)	(0.008)	(0.008)	(0.008)	(0.008)
	355	*1154*	*105*	*145*	*156*
26	-0.073***	-0.078***	-0.081***	-0.086***	-0.092***
	(0.008)	(0.008)	(0.008)	(0.008)	(0.007)
	291	*1069*	*106*	*190*	*214*
27	-0.051***	-0.056***	-0.059***	-0.064***	-0.070***
	(0.007)	(0.007)	(0.007)	(0.007)	(0.007)
	288	*952*	*149*	*231*	*325*
28	-0.030***	-0.035***	-0.038***	-0.043***	-0.048***
	(0.007)	(0.007)	(0.008)	(0.009)	(0.009)
	268	*822*	*130*	*250*	*349*
29	-0.008	-0.013	-0.016*	-0.021**	-0.026***
	(0.009)	(0.009)	(0.009)	(0.009)	(0.009)
	227	*757*	*122*	*243*	*365*
30	0.014	0.009	0.006	0.001	-0.005
	(0.009)	(0.008)	(0.008)	(0.008)	(0.008)
	194	*606*	*104*	*197*	*329*
31	0.036***	0.030***	0.028***	0.022**	0.017*
	(0.008)	(0.008)	(0.008)	(0.009)	(0.008)
	179	*461*	*82*	*165*	*271*
32	0.057***	0.052***	0.049***	0.044***	0.039***
	(0.008)	(0.008)	(0.008)	(0.008)	(0.008)
	111	*377*	*61*	*124*	*219*
33	0.079***	0.074***	0.071***	0.066***	0.060***
	(0.008)	(0.008)	(0.007)	(0.007)	(0.007)
	87	*281*	*58*	*102*	*167*
34	0.101***	0.095***	0.093***	0.087***	0.082***
	(0.007)	(0.007)	(0.007)	(0.007)	(0.007)
	74	*253*	*28*	*73*	*114*

Notes: *p<0.1; **p<0.05; ***p<0.01. Bootstraped standard errors in parentheses and number of observations in each cell in italics.

Table 10: The effect of having a third child on labor market duration for mothers with $Q_1 < DW \leq Q_2$ for the restricted sample of mothers born 1923–1935

Age at second birth	Years of education 9	11	12	14	16
22	-0.147*** (0.009) *188*	-0.151*** (0.009) *772*	-0.153*** (0.009) *40*	-0.157*** (0.009) *48*	-0.161*** (0.009) *44*
23	-0.129*** (0.009) *262*	-0.133*** (0.009) *891*	-0.135*** (0.009) *60*	-0.139*** (0.009) *80*	-0.143*** (0.008) *67*
24	-0.111*** (0.008) *287*	-0.115*** (0.008) *995*	-0.117*** (0.008) *94*	-0.121*** (0.008) *97*	-0.125*** (0.008) *115*
25	-0.093*** (0.008) *299*	-0.097*** (0.008) *977*	-0.099*** (0.008) *95*	-0.103*** (0.008) *141*	-0.107*** (0.008) *182*
26	-0.075*** (0.008) *349*	-0.079*** (0.008) *1016*	-0.081*** (0.008) *128*	-0.085*** (0.008) *196*	-0.089*** (0.007) *273*
27	-0.057*** (0.007) *265*	-0.061*** (0.007) *950*	-0.063*** (0.007) *143*	-0.067*** (0.007) *273*	-0.071*** (0.007) *374*
28	-0.039*** (0.007) *233*	-0.043*** (0.007) *872*	-0.045*** (0.007) *136*	-0.049*** (0.009) *270*	-0.053*** (0.009) *452*
29	-0.021** (0.009) *207*	-0.025*** (0.009) *722*	-0.027*** (0.009) *122*	-0.031*** (0.009) *281*	-0.035*** (0.009) *466*
30	-0.003 (0.009) *191*	-0.007 (0.009) *588*	-0.009 (0.008) *98*	-0.013 (0.008) *246*	-0.017** (0.008) *442*
31	0.015* (0.008) *120*	0.011 (0.008) *479*	0.009 (0.008) *96*	0.005 (0.009) *194*	0.001 (0.008) *286*
32	0.033*** (0.008) *110*	0.029*** (0.008) *357*	0.027*** (0.008) *67*	0.023*** (0.008) *177*	0.019** (0.008) *264*
33	0.051*** (0.008) *86*	0.047*** (0.008) *304*	0.045*** (0.007) *43*	0.041*** (0.007) *132*	0.037*** (0.007) *139*
34	0.069*** (0.007) *71*	0.065*** (0.007) *192*	0.063*** (0.007) *22*	0.059*** (0.007) *69*	0.055*** (0.007) *113*

Notes: *p<0.1; **p<0.05; ***p<0.01. Bootstraped standard errors in parentheses and number of observations in each cell in italics.

Table 11: The effect of having a third child on labor market duration for mothers with $Q_2 < DW \leq Q_3$ for the restricted sample of mothers born 1923–1935

	Years of education				
Age at second birth	9	11	12	14	16
22	-0.130***	-0.133***	-0.135***	-0.138***	-0.141***
	(0.009)	(0.009)	(0.009)	(0.009)	(0.009)
	178	*795*	*51*	*52*	*38*
23	-0.115***	-0.118***	-0.120***	-0.123***	-0.126***
	(0.009)	(0.009)	(0.009)	(0.009)	(0.009)
	294	*1011*	*63*	*64*	*58*
24	-0.100***	-0.103***	-0.105***	-0.108***	-0.111***
	(0.008)	(0.008)	(0.008)	(0.008)	(0.008)
	297	*1202*	*102*	*112*	*115*
25	-0.085***	-0.088***	-0.090***	-0.093***	-0.096***
	(0.009)	(0.008)	(0.008)	(0.008)	(0.008)
	352	*1249*	*121*	*150*	*163*
26	-0.070***	-0.073***	-0.075***	-0.078***	-0.081***
	(0.008)	(0.008)	(0.008)	(0.008)	(0.007)
	386	*1294*	*147*	*220*	*282*
27	-0.055***	-0.058***	-0.060***	-0.063***	-0.066***
	(0.007)	(0.007)	(0.007)	(0.007)	(0.007)
	355	*1140*	*191*	*288*	*392*
28	-0.040***	-0.044***	-0.045***	-0.048***	-0.051***
	(0.007)	(0.007)	(0.007)	(0.009)	(0.009)
	303	*1042*	*187*	*300*	*384*
29	-0.026***	-0.029***	-0.030***	-0.033***	-0.036***
	(0.009)	(0.009)	(0.009)	(0.009)	(0.009)
	302	*896*	*145*	*290*	*431*
30	-0.011	-0.014*	-0.015**	-0.018**	-0.022***
	(0.009)	(0.009)	(0.009)	(0.009)	(0.008)
	216	*724*	*112*	*222*	*388*
31	0.004	0.001	-0.000	-0.004	-0.007
	(0.008)	(0.008)	(0.009)	(0.009)	(0.008)
	152	*507*	*92*	*182*	*303*
32	0.019**	0.016**	0.014*	0.011	0.008
	(0.008)	(0.008)	(0.008)	(0.008)	(0.008)
	114	*394*	*62*	*137*	*199*
33	0.034***	0.031***	0.029***	0.026***	0.023***
	(0.008)	(0.008)	(0.007)	(0.007)	(0.007)
	57	*252*	*48*	*72*	*134*
34	0.049***	0.046***	0.044***	0.041***	0.038***
	(0.007)	(0.007)	(0.007)	(0.007)	(0.008)
	49	*136*	*28*	*59*	*75*

Notes: *p<0.1; **p<0.05; ***p<0.01. Bootstraped standard errors in parentheses and number of observations in each cell in italics.

Table 12: The effect of having a third child on labor market duration for mothers with $Q_3 < DW$ for the restricted sample of mothers born 1923–1935

Age at second birth	Years of education: 9	11	12	14	16
22	-0.103***	-0.105***	-0.106***	-0.109***	-0.112***
	(0.010)	(0.009)	(0.009)	(0.009)	(0.009)
	222	*849*	*50*	*54*	*48*
23	-0.092***	-0.095***	-0.096***	-0.099***	-0.101***
	(0.009)	(0.009)	(0.009)	(0.009)	(0.009)
	271	*1085*	*71*	*95*	*55*
24	-0.082***	-0.085***	-0.086***	-0.088***	-0.091***
	(0.009)	(0.009)	(0.009)	(0.009)	(0.009)
	306	*1177*	*86*	*104*	*72*
25	-0.072***	-0.074***	-0.076***	-0.078***	-0.081***
	(0.009)	(0.008)	(0.008)	(0.008)	(0.008)
	343	*1132*	*98*	*123*	*122*
26	-0.062***	-0.064***	-0.065***	-0.068***	-0.070***
	(0.008)	(0.008)	(0.008)	(0.008)	(0.008)
	330	*1103*	*122*	*125*	*161*
27	-0.051***	-0.054***	-0.055***	-0.058***	-0.060***
	(0.007)	(0.007)	(0.007)	(0.007)	(0.007)
	269	*970*	*109*	*139*	*198*
28	-0.041***	-0.044***	-0.045***	-0.047***	-0.050***
	(0.007)	(0.008)	(0.008)	(0.010)	(0.010)
	257	*777*	*98*	*138*	*205*
29	-0.031***	-0.033***	-0.035***	-0.037***	-0.040***
	(0.010)	(0.009)	(0.009)	(0.010)	(0.010)
	221	*636*	*84*	*132*	*181*
30	-0.021***	-0.023***	-0.024***	-0.027***	-0.029***
	(0.010)	(0.009)	(0.009)	(0.009)	(0.009)
	168	*436*	*73*	*93*	*153*
31	-0.010	-0.013*	-0.014**	-0.016**	-0.019**
	(0.009)	(0.009)	(0.009)	(0.009)	(0.008)
	116	*276*	*41*	*83*	*121*
32	-0.000	-0.002	-0.004	-0.006	-0.009
	(0.008)	(0.008)	(0.008)	(0.008)	(0.008)
	63	*175*	*23*	*55*	*67*
33	0.010	0.008	0.006	0.004	0.002
	(0.008)	(0.008)	(0.008)	(0.008)	(0.007)
	35	*114*	*17*	*30*	*44*
34	0.020***	0.018**	0.017**	0.014*	0.012
	(0.007)	(0.007)	(0.007)	(0.008)	(0.008)
	20	*49*	*9*	*16*	*17*

Notes: *p<0.1; **p<0.05; ***p<0.01. Bootstraped standard errors in parentheses and number of observations in each cell in italics.

Given age and education, also short waiting times betwen the second and third birth in contrast to long can be taken as a measure of labor market attachment. That is, women with a high labor market attachment, all else equal, return to work faster in between child births than those with a low labor market attachment. This line of argument is supported by the negative effects of waiting time. It is interesting to note that the results on waiting time, waiting time interacted with age at second birth, and waiting time interacted with education are reversed from what was found when not controlling for endogeneity (see column 2 in tables 3 and 8). We believe the reason for this is that without controlling for endogeneity, the two first variables cannot be seen as proxies of labor market attachment as they influence the choice of having a third child.

The result of a less negative or even a positive effect on long-term labor supply for low educated women may at a first glance seem implausible. However, such an effect could stem from the increased economic costs of having three children in contrast with two. Women with lower education may have stronger financial restrictions compared to highly educated mothers, which restrict them from leaving the labor market early.

We also find that the range of the effect magnitude at different ages of second child birth is monotonically decreasing with child spacing. This results most likely stems from younger mothers with short waiting times being more likely to have more than three children, compared to those with long waiting times. Due to decreasing fertility over age, the same problem is not so pertinent to the mothers who are older at second birth. In figure 2 in Appendix B, we present the tabulation of the probability of having a fourth child contingent on W and the age at second birth. The results from this tabulation verify our speculation.

6 Conclusion

The paper proposes a general estimation framework for studying the long-term impact of fertility on female labor market work duration. Although the framework does not provide direct estimates of the effect of having an extra child on life-time income, it provides a lower bound under the reasonable assumption that the effect of fertility on wages is non-positive. The identification of the effect of fertility is difficult because fertility and career decisions most likely are jointly determined. For this reason, some form of exogenous variation in fertility is useful. Here we use the fact that parents have preferences for a mixed sex sibling composition and use the ‘same-sex’ instrument. Consequently, in the empirical application, we study the impact of having a third child on female labor market work duration.

One problem neglected in most studies using a quasi-experimental design is that the timing of fertility is chosen. We show that the genuine treatment is waiting time to birth rather than birth per se. As the same sex instrument affects fertility, it also affects the spacing between the second and third child, which enables us to also estimate the effect of child spacing on female labor market work duration.

In the estimation we use data from Swedish administrative registers, providing total coverage for those who were born in 1923 or later. We sample mothers to second-born children from the register and then – unless not censored due to end of study – we can follow them until retirement. We restrict the main analysis to mothers born in 1923 and 1947. However as there were large changes in both the tax and social insurance system in the 1970's we also perform an analysis to mothers born between 1923-1935. The women born in 1935 are 35 years in 1970.

The results from both analyses is that having a third child will in general reduce the labor market work duration. The magnitude of the effect depends to a large extent on the age of the women when giving second birth, but also on the waiting time to the third child and the education level. The results from the two analyses suggest that for a mother with average education, average age at second birth and the shortest birth spacing, the reduction in the labor market duration is around 5.2 percent and 3.2 percent, respectively (or 1.8 years and 1.1 years, respectively). For highly educated mothers who had her second child early, the reduction in labor market work duration can even be above 30 percent. For mothers who had their second child late we even find positive labor market work duration effects. This results is more pronounced for mothers who, in addition to giving second birth late, have a short spacing between the second and third child and low education. We believe these effect differences stem from a higher labor market attachment for the women having their second child late compared to younger mothers. This interpretation is strengthened by the less negative effect of short waiting times in contrast to long. We believe the reason for the lower or even positive effects for the less educated mothers might have to do with financial restrictions. Having an additional child is costly and might force low-educated mothers to prolong their time on the labor market.

We have shown that the labor market consequences of an extra child (from two to three) are not equally shared among mothers with three children. Women having children early on in their career may be more affected than women who are well established on the labor market. From a policy perspective it becomes clear that a social insurance system with maternity leave based on income protection, which stimulates labor market attachment before child bearing, would hence moderate the long term effects on the labor market. In this respect the less pronounced negative effect when removing women affected by the reforms in the 1970's (i.e., those born after 1935), was a surprise. Given that the reform implemented in 1970's encouraged labor supply of women we would have expected the opposite result. However, the effects on third child could be different than the effects on parenthood per se. Working women with more than two children in the older cohorts could very well be a positively selected population in terms of labor market attachment, compared to later cohorts. Accordingly, this would attenuate the effect from a third child for this group. One interpretation of the results from this study is hence that as long as mothers take the main responsibility for the rearing and caring of the children, there will still be a penalty in terms of lower income and wages for the mothers but also in labor market work duration. The implication is hence that current child-friendly policies aiming at an increased female labor supply may not be the only way forward for gender equality on the labor market.

As this is the first study on this outcome we cannot compare the magnitude of the effects to other countries with other institutions. Therefore, replication using data under other institutional settings than the Swedish would be of great interest as well as to repeat this study (when data exists) on later cohorts with a more extensive social insurance system.

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Appendix A: Generalized residuals

To derive the generalized residuals, we use the well-known moment formulas for truncated $N(0,1)$: for $\kappa \sim N(0,1)$,

$$E(\kappa|\kappa > c) = \frac{\phi(-c)}{\Phi(-c)},\ E(\kappa^2|\kappa > c) = 1 + c\frac{\phi(-c)}{\Phi(-c)},\ E(\kappa^3|\kappa > c) = \frac{\phi(-c)}{\Phi(-c)}(2+c^2).$$

Let $A \equiv \ln(45 - T_2 + T_b) - X'\alpha_x - Z'\alpha_z$. Then the generalized residuals come from

$$\begin{aligned}
E(\varepsilon|\varepsilon &> A) = \sigma_\varepsilon E\left(\frac{\varepsilon}{\sigma_\varepsilon}|\frac{\varepsilon}{\sigma_\varepsilon} > \frac{A}{\sigma_\varepsilon}\right) = \sigma_\varepsilon E\left(\kappa|\kappa > \frac{A}{\sigma_\varepsilon}\right) = \sigma_\varepsilon \frac{\phi(-A/\sigma_\varepsilon)}{\Phi(-A/\sigma_\varepsilon)};\\
E(\varepsilon^2|\varepsilon &> A) = \sigma_\varepsilon^2 E\left(\frac{\varepsilon^2}{\sigma_\varepsilon^2}|\frac{\varepsilon}{\sigma_\varepsilon} > \frac{A}{\sigma_\varepsilon}\right) = \sigma_\varepsilon^2 E\left(\kappa^2|\kappa > \frac{A}{\sigma_\varepsilon}\right) = \sigma_\varepsilon^2\left\{1 + \frac{A}{\sigma_\varepsilon}\frac{\phi(-A/\sigma_\varepsilon)}{\Phi(-A/\sigma_\varepsilon)}\right\};\\
E(\varepsilon^3|\varepsilon &> A) = \sigma_\varepsilon^3 E\left(\frac{\varepsilon^3}{\sigma_\varepsilon^3}|\frac{\varepsilon}{\sigma_\varepsilon} > \frac{A}{\sigma_\varepsilon}\right) = \sigma_\varepsilon^3 E\left(\kappa^3|\kappa > \frac{A}{\sigma_\varepsilon}\right) = \sigma_\varepsilon^3 \frac{\phi(-A/\sigma_\varepsilon)}{\Phi(-A/\sigma_\varepsilon)}\left\{2 + \left(\frac{A}{\sigma_\varepsilon}\right)^2\right\}.
\end{aligned}$$

Appendix B

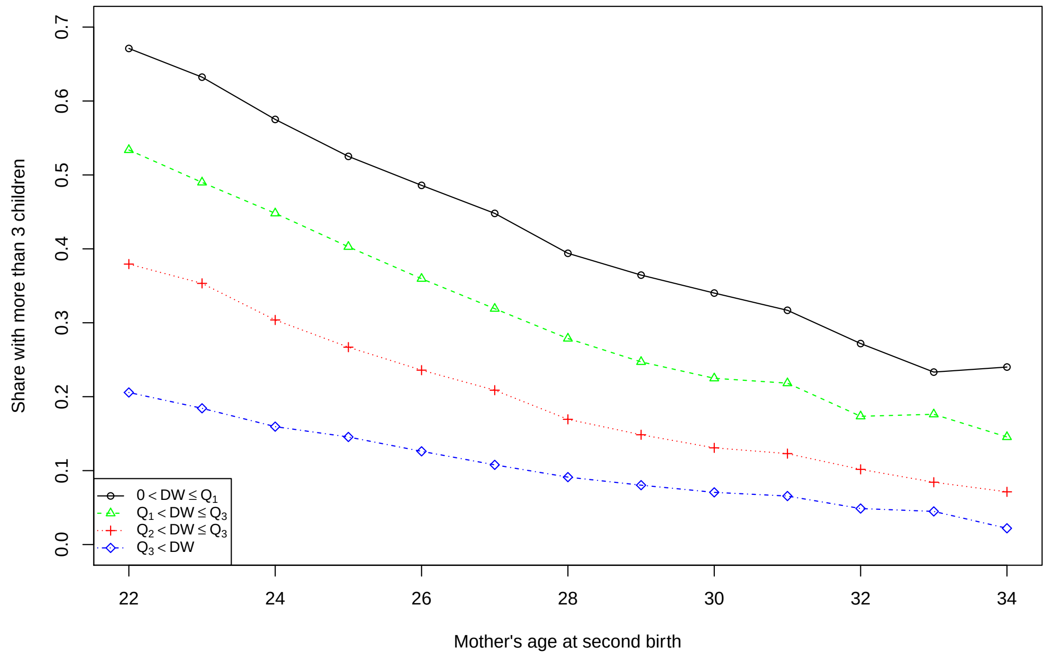

Figure 2: Shares of women with four or more children by age at second birth and child spacing between second and third birth.