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Effects of the timing of births on women's earnings – evidence from a natural experiment

Arizo Karimi

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Effects of the timing of births on women's earnings - evidence from a natural experiment[a]

by

Arizo Karimi[b]

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Abstract
This paper studies the causal effect of the timing of first birth on highly educated women's career outcomes using exogenous variation in first birth timing induced by the occurrence of pregnancy loss before first birth. Contrasting previous findings, my results suggest that a one-year delay has a significantly negative effect on both income and wages. The negative effects might partly be explained by child spacing; motherhood delay induces women to have the second child more closely spaced (but not fewer or more children altogether), and consequently to have a potentially longer consecutive parental leave, or more frequent transitions in and out of the labor market. The same findings hold true when I employ an individual-fixed effects estimator based on panel data, from which the results suggest a larger slope decline in the wage profile post birth for "late" mothers. The hypothesis that short birth intervals may be detrimental for career outcomes is then tested by analyzing the impact of spacing births, using miscarriages between the first and second births as an instrument for birth spacing. The results suggest that a longer birth spacing indeed has positive long-run effects on income and wage rates.

Keywords: First-birth timing, child spacing, female wages, lifetime earnings
JEL-codes: J31, J13

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1 Introduction

There is a vast economics literature that addresses the questions of how fertility is related to women's labor supply, income and wages. It is well established that childbearing reduces women's subsequent labor supply and income (see e.g. Bronars & Grogger 1994, Angrist & Evans 1998, Jacobsen et al. 1999, Vere 2011), and career interruptions due to childbearing might also imply lower subsequent wage rates through the foregone investments in human capital and possible skill depreciation. Thus, work interruptions - and the career costs that they entail - constitute a major component of the opportunity costs of children. Recently, researchers have devoted increasing attention towards the question of whether the timing of parenthood can affect the magnitude of such costs. Partly, this interest has been spurred from the empirical observation that the age at first birth is positively associated with labor market outcomes for women (see e.g. Chandler et al. 1994, Hofferth 1984). In addition, many developed countries have witnessed an increase in the age at first birth while simultaneously witnessing improvements in women's labor market status. Indeed, previous studies suggest that postponing motherhood positively affects mothers' subsequent earnings (Miller 2011, Taniguchi 1999, Amuedo-Dorantes & Kimmel 2005, Troske & Voicu 2012, Bratti & Cavalli 2014).[1] However, there is to date limited knowledge about the potential mechanisms through which postponing motherhood positively affects income and wage rates. For instance, potential mediating channels include the total number of children as well as the tempo of subsequent fertility, as both factors are likely to have implications for the duration of work interruptions. Thus, to assess the total effect of first birth timing, it is potentially important to take into consideration and analyze the impacts of first birth timing on subsequent fertility outcomes.

The aim of this paper is to estimate the impact of first birth timing on the income and wage rates over the careers of highly educated women - the group most often observed to postpone motherhood. In addition, I estimate the effect of first birth timing on parental leave usage, the total number of children born to a woman, as well as on the time interval

[1] Related to studies on the relationship between age at first birth are studies on the impacts of teenage childbearing, with some studies finding negative effects of early childbearing (e.g Klepinger et al. 1999), some studies finding modest effects (Holmlund 2005, Ashcraft et al. 2013) and some positive effects of early childbearing (Ribar 1996, Hotz et al. 1997, 2005).

to the subsequent birth (i.e., on child spacing), all of which are likely to be important determinants for long-run labor market outcomes.

An underlying challenge in isolating a causal relationship, however, is the potential endogeneity of the timing of entry into motherhood with respect to labor market outcomes. One possible source of endogeneity in this context is that individuals are likely to exhibit unobserved heterogeneity in tastes or motivation that affects both fertility and career choices. Moreover, it seems likely that fertility timing choices respond to anticipated career outcomes, or that choices about fertility and careers are jointly determined. I exploit a source of exogenous variation in birth timing induced by miscarriage before first birth, used in previous studies by Miller (2011), Hotz et al. (1997, 2005), Ashcraft et al. (2013), Bratti & Cavalli (2014), Herr (2007). Pregnancy losses are naturally occurring fertility shocks that delay time to birth and can thus be used as an instrument for first birth timing. As an alternative empirical strategy, I also employ an individual-fixed effects estimator using panel data. The effect of first birth timing is thus estimated under two different sets of identifying assumptions and the analyses provided allow comparisons of estimates relying on different sources of variation. To this end, I use a combination of different Swedish registers with individual level information on income, wage rates, background characteristics, parental leave usage, sickness absence, and fertility histories. Individual level data on miscarriages are drawn from hospital registers, which include detailed information on the causes of admissions in the form of medical diagnoses classified according to the International Classification standard for Diseases (ICD).[2]

While using the same strategy employed by Miller (2011) and Hotz et al. (2005), the present paper contributes to the existing findings in several ways. First, most of the previous evidence is based on data from the United States. Sweden is an interesting case to study in this context since the institutional setting differs considerably from that of the United States. In particular, family policies are universal and generous, including job-protected parental leave and governmentally funded wage replacement. Thus, while postponing fertility may be important in countries without state mandated job-protection

[2]One advantage with using the National Patient Register over relying on survey data - which has been the main type of data source used in previous studies - is that I avoid potential misreporting of abortions as miscarriages, which might be likely if the re is a social stigma associated with abortions.

to, e.g., gain a suitable job-match that allows a non-disruptive career with childbearing, it is most likely not needed in Sweden where job-protected leave is the default. In general, I expect the benefits to delayed motherhood to be smaller in Sweden since family policies, in particular job-protection during parental leave, should lessen the tradeoff between family and market work.

Focusing on women who first finish college, then enter the labor market and subsequently become mothers, I estimate the effect of having one extra year of labor market experience before entry into motherhood on the total income earned over the first 20 years of the career, and on the average wage rate over the career. The results obtained from the OLS estimator suggest a positive relationship between the the number of years of pre-motherhood work experience before first birth (i.e., birth timing) and the net present value of earnings of about 4 percent. However, when instrumenting first birth timing with the occurrence of miscarriage, I find that postponing motherhood has a significantly negative effect on career earnings and on the average career wage of about 15 and 5 percent, respectively. Estimating the effect of postponing first birth on the *yearly* income post birth (with 2SLS) shows that earnings are lower in the first four years after birth compared to what they would have been had the first birth not been delayed. Thus, not only is the benefit to motherhood delay smaller in Sweden, motherhood delay in fact appears to have adverse career consequences. The negative effects of postponing motherhood among adult, highly educated mothers contrasts results from previous studies (e.g. Miller 2011, Bratti & Cavalli 2014).[3] Studying sickness absence before and after childbearing/miscarriage does not lend support to the conjecture that the negative effects are driven by a violation of the exclusion restriction through adverse effects of miscarriage on health. In order to shed some light on the potential mechanisms driving these effects, a second contribution of this paper is to study the effects of motherhood postponement on completed fertility and child spacing. The results from this analysis show that postponing first birth does not affect the total number of children, but instead accelerates the time to next birth. This could

[3] However, using a dynamic treatment approach to study the effect of having a first child at a certain age against the alternative of delaying childbearing at that age on subsequent employment, Fitzenberger et al. (2013) find no evidence supporting the hypothesis that delayed childbearing reduces the negative employment effects found of childbearing.

imply being absent from the labor market for a longer portion of a potentially critical period of career build-up. Analyses of the parental leave usage response indeed suggests that parental leave usage increases in the immediate years following birth relative to what would have occurred had the first birth not been delayed. The time pattern exhibited by the parental leave response is consistent with the time pattern exhibited by subsequent fertility responses: postponing first birth induces an increase in the probability to give birth to a subsequent child two years after first birth, while inducing negative effects on the probability to give birth to a subsequent child in later years following first birth: thus, a shift in the timing of second birth to occur more closely after the first birth. Corroborating the results from the instrumental variables analysis, an individual-fixed effects estimator suggests that the slope decline in wages post birth is larger for "late" mothers compared to "earlier" mothers.

A third contribution of this paper stems from a more explicit test of the hypothesis that the shorter time interval between births, caused by postponing the entry into motherhood, is the driving mechanism for the negative effects of delayed motherhood on earnings. Specifically, I estimate the causal effect of birth spacing using the exogenous variation in spacing induced by pregnancy loss between the first and second live birth. The results from this analysis are in line with the conjecture that two closely spaced births may be detrimental for women's careers: a longer birth interval is found to have sizeable positive effects on income and wage rates lasting even 15 years after the second child is born.

The results provided in this paper thus imply that the same number of work interruptions over the career may have different consequences for labor market attachment and wage rates depending on how closely these interruptions are spaced. This raises questions about depreciation of general human capital being the main mechanism through which motherhood affects wages. In addition to general human capital, connectivity to the labor market more generally could affect firm-specific human capital as well as signals about work commitment to the employer, or even the loss of a job-specific network. Moreover, from a policy point of view, the question of whether the timing of parenthood matters for labor market outcomes is particularly interesting since the age at first birth has been observed to be responsive to policy changes. For instance, Björklund (2006) reports that the family policies introduced in Sweden between 1960 and 1980, in which benefits were tied to previous labor earnings, increased women's age at first birth.[4] With regards to the spacing of births, the average time interval between children in Sweden has decreased over time, in part due to the introduction of the so called "speed premium" rules in the parental leave system, exempting mothers from re-establishing eligibility to paid leave for a subsequent child by re-entering the labor market after first birth, granted that the next child is born sufficiently close to the preceding child. Thus, family policies may have unintended consequences for the timing of fertility, making it relevant to understand the impacts of fertility timing on labor market outcomes. This becomes especially important if mothers' earnings matter for child well-being, and thus if spacing births have effects on the next generation through the households' financial resources.

2 Identification strategy

The objective of this paper is to estimate the effect of first birth timing on women's career outcomes. I measure first birth timing as the number of years elapsed between labor market entry and first birth - as opposed to the age at first birth. This implies that fertility timing here can be thought of as a more direct measure of potential (pre-natal) experience, compared to experience being proxied by age.[5] Setting up the problem in a potential-outcomes framework, let Y denote the labor market outcome of interest and let T denote first birth timing, measured as the number of years elapsed between labor market entry and the birth year of the first child (i.e., pre-birth labor market experience). We are worried that the regressor of interest, T, might be endogenous to labor market outcomes, Y, due to unobserved heterogeneity in tastes or motivation that affects both fertility and work choices. To address the potential endogeneity issues, I make use of the exogenous source of variation in first birth timing induced by the event of miscarriage before first birth, the

[4] Further, in an overview on the effects of family policies in industrialized countries, Gauthier (2007) reports that some studies suggest that the effect of policies tend to be on the timing of births rather than on completed fertility.

[5] As suggested by Herr (2011), this definition of birth timing might be more appropriate than age at first birth, based on her findings that the latter tends to underestimate the return to motherhood delay for women who remain childless at labor market entry, and obscure the negative return to delay to a first birth after labor market entry for all but college graduates.

incidence of which extends time to motherhood (Miller 2011, Hotz et al. 1997, 2005). Let the binary variable Z indicate first pregnancy ending in miscarriage. Then, let T_1 denote the first birth timing for an individual with $Z = 1$ and let T_0 denote the timing for an individual with $Z = 0$. Moreover, we can consider T a treatment with variable treatment intensity, taking on the values $j = 0, 1, 2, ..., J$. Suppose that each individual would earn Y_j if she waited j years between entering the labor market and entering motherhood, for $j = 0, 1, 2, ..., J$. While a full set of Y_j is well defined for each individual, only one is ever observed. The goal is to attain information about the distribution of $Y_j - Y_{j-1}$, which is the causal effect of the first career interruption due to childbearing occurring in the j:th year.

For Z to be a valid instrument, the first identifying assumption that needs to hold is that Z is independent of all potential outcomes and potential treatment intensities, i.e., that $T_0, T_1, Y_1, ...Y_J$ are jointly independent of Z. Independence alone is not always sufficient to estimate a meaningful average treatment effect, since it is theoretically possible to have a situation where the treatment effect is positive for everyone, but the sizes of the groups of compliers and defiers are such that the average difference in outcomes is zero or even negative. To get around this problem, a second identifying assumption needed is monotonicity: With probability 1, $T_1 - T_0 \geq 0$ for each person. Given independence, monotonicity and the assumption that $Pr(T_1 \geq j > T_0)$, i.e. that there exists a First-stage relationship, for at least one J, Angrist & Imbens (1995) show that

$$LATE = \frac{E(Y|Z=1) - E(Y|Z=0)}{E(T|Z=1) - E(T|Z=0)}$$

$$= \sum_{j=1}^{J} \omega_j E[Y_j - Y_{j-1} | T_1 \geq j > T_0] \equiv \beta$$

where

$$\omega_j = \frac{P(T_1 \geq j > T_0)}{\sum_{i=1}^{J} Pr(T_1 \geq i > T_0)}$$

with $0 \leq \omega_j \leq 1$ and $\sum_{j=1}^{J} \omega_j = 1$, so that β is a weighted average of a per-unit treatment effect. Angrist & Imbens (1995) refer to β as the average causal response (ACR). This parameter captures a weighted average of causal responses to a unit change in treatment, for those whose treatment status is affected by the instrument. The weight attached to the average of $Y_j - Y_{j-1}$ is proportional to the number of people who, because of the instrument, change their treatment intensity from less than j units to j or more units. This proportion is $Pr(T_1 \geq j > T_0)$; the proportion who, by the event of experiencing a miscarriage, are induced to delay motherhood.

As shown in Angrist & Imbens (1995), for a multi-valued treatment ($J > 1$), the monotonicity assumption has the testable implication that the cumulative distribution function (CDF) of T given $Z = 1$ and the CDF of T given $Z = 0$ should not cross.[6] Although there is no direct reason to be worried that the monotonicity assumption does not hold in the case of miscarriages (there can be no defiers by construction because a miscarriage always delays births), we can plot the empirical CDF:s to gain knowledge about the weighting function of the ACR. The CDF:s for birth timing, by the value of the instruments, are graphed in the upper panel of *Figure A1* in the Appendix, along with the best fitted normal model superimposed over the sample CDF. The figure shows that for mothers who experienced a miscarriage, the CDF lies below the CDF for women who did not experience a miscarriage until timing, i.e., T, equals 10. After year 10, the CDF:s cross, and the CDF for women with miscarriages lies above the CDF for women with no miscarriage. This evidence is in support of the monotonicity assumption for those mothers who wait at most 10 years after entering the labor market until they have their first child. One possible explanation is that for some women who wait a long time, miscarriages might be indicative of them trying harder to get pregnant. In the analyses, I will perform the estimations for the sub-samples of mothers with first birth timing less than 11 years.

Furthermore, the weighting function of the ACR for estimates based on comparisons between women who do and do not experience a miscarriage is the difference between the CDF:s normalized to sum to one. This difference is plotted in the lower panel of *Figure A1* and shows that the group contributing most to the estimates of the ACR based on the event of miscarriage are those with 2-3 years elapsed between entering the labor

[6] If $T_1 \geq T_0$ with probability 1, then $Pr(T_1 \geq j) \geq Pr(T_0 \geq j)$ for all j, which implies $Pr(T \geq j|Z = 1) \geq Pr(T \geq j|Z = 0)$ or $F_T(j|Z = 1) \geq F_T(j|Z = 0)$, where F_T is the CDF of T (Angrist & Imbens 1995).

market and having a first child. At most 9 percent of the sample was induced by having a miscarriage to have their first child in career year 3, but smaller fractions were induced to have their first child at later career points.

2.1 Threats to identification

While the existence of a First-stage relationship can be directly addressed, the independence condition cannot be formally tested. One potential concern with instrumenting birth timing with miscarriage is that the health of mothers who miscarry is, on average, worse compared to women who do not. These health limitations in turn would lead women to have lower wages. Another concern is that miscarriages might cause psychological distress and therefore directly affect labor market outcomes, violating the exclusion restriction. This critique against using miscarriage as a source of variation in birth timing is lifted by e.g. Wilde et al. (2010), who in addition also worry that behavioral characteristics differ between women who miscarry and women who do not. For instance, some evidence suggest that miscarriage risk is associated with risky behaviors such as regular or high alcohol consumption, tobacco or drug use during pregnancy (see e.g. Garcia-Enguidanos et al. 2002, Maconochie et al. 2007, for overviews). Garcia-Enguidanos et al. (2002), however, argue that while many risk factors have been suggested in the medical literature, there are only two factors recognized by the studies included in their review, which are uterine malformations and chromosomal rearrangements. Moreover, miscarriage is a frequently occurring fertility shock; Regan & Rai (2000) review the medical literature and state that sporadic miscarriage is the most common complication of pregnancy, and one in four of all women who become pregnant will experience pregnancy loss. Moreover, the vast majority of pregnancy losses are early, occurring well before 12 weeks of gestation, with sporadic miscarriage after this time complicating no more than 1-2 percent of pregnancies (Regan & Rai 2000).

To investigate whether there are health differences between women who miscarry and women who do not in my sample, I make use of detailed individual level data from the National Patient Register (NPR), which covers the universe of all hospitalizations in Sweden between 1987 and 2005. The NPR is an inpatient care record that includes medical information in the form of the diagnosis associated with each hospital visit, classified according to the International Classification Standard for Diseases (ICD).[7] Using these data, I study whether there are any differences in pre-motherhood incidence of hospitalizations between mothers with $Z = 1$ and mothers with $Z = 0$, i.e., between mothers who did and did not experience a miscarriage before first birth. The results from this analysis are presented in *Figure 2.1*, where the upper graph plots the differences in the average number of hospital visits for different diagnoses during a time period of 4 years before first birth (birth year -4 to birth year -1) with the corresponding 95 percent confidence intervals. A first thing to note is that there are, if any, very small differences in the frequency of hospitalizations between women who miscarry and women who do not for any of the diagnoses; in fact, most estimates lie on the vertical zero line.[8] Nevertheless, some diagnoses are shown to be more prevalent among women who miscarried. For instance, women who experienced a miscarriage had somewhat higher frequency of pre-natal hospital visits due tumors and neoplasm diseases and respiratory and endocrine diseases. These differences are, however, very small. The largest difference is found in the average number of hospital visits associated with diseases of the genitourinary system, which are more prevalent among women who later experienced a miscarriage. There are no indications of differences between the groups concerning hospitalizations associated with risky behaviors such as alcohol or substance abuse. This does probably not imply that alcohol or drug use are not risk factors for miscarriage; admissions due to these causes are very uncommon for the studied sample, making it difficult to draw clear conclusions on the extent to which risky behavior differ between women who do and do not experience miscarriage. Given the statistically significant differences in hospital admissions for some of the diagnoses categories, it is important to control for for health. Thus, in all analyses I control for the number of pre-natal hospitalizations including the number of hospital visits associated with each of the diagnosis category depicted in *Figure 2.1*.

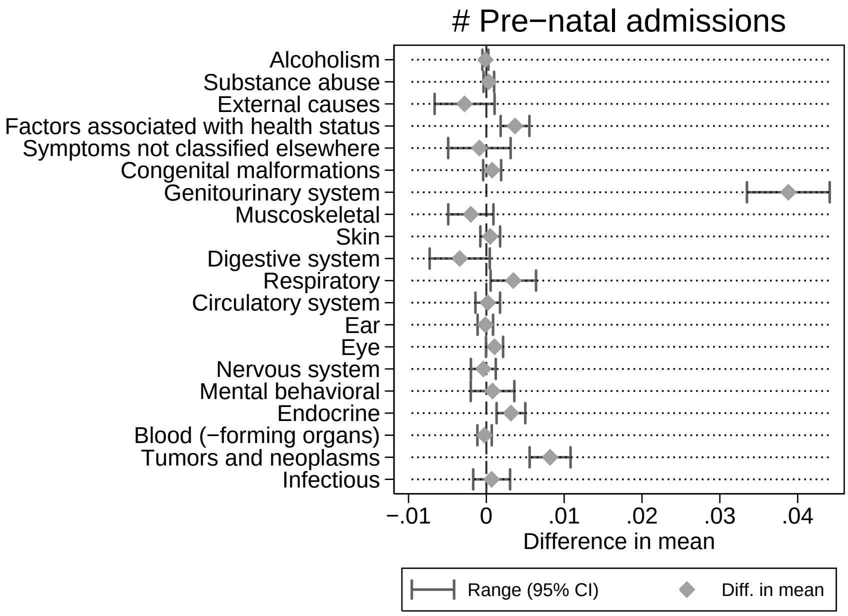
The lower graph in *Figure 2.1* is analogous to the upper graph, but presents differences in average personal characteristics as well as in the first birth timing. As seen, women who

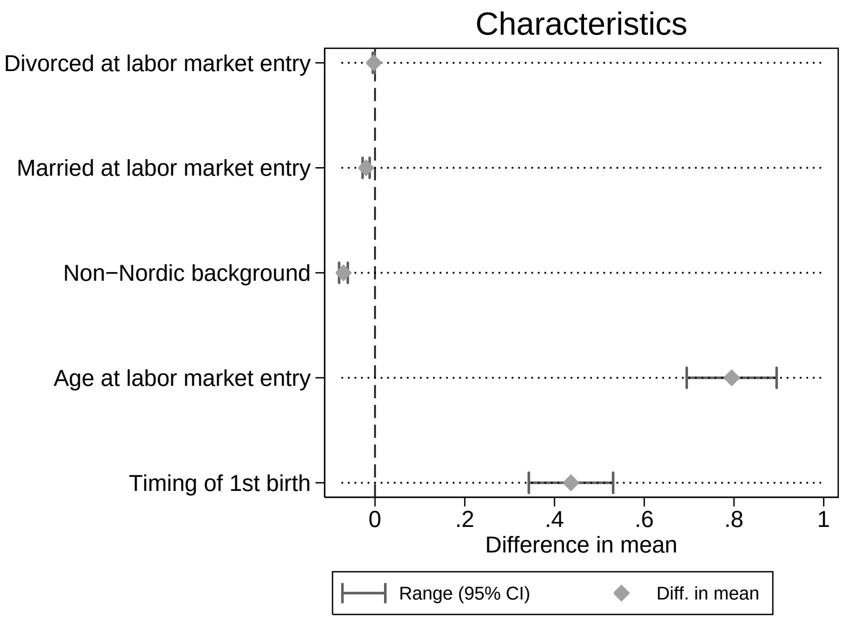
[7]The NPR is also the data source I use for identifying miscarriage events; a more detailed description of the data follows in the Data section.

[8]Hospitalizations due to pregnancy-related reasons, including miscarriage, are not included in this analysis.

experienced a miscarriage had a delayed childbearing by on average 6 months (this is the "raw" first-stage estimate). There are hardly any differences in marital status at the time of labor market entrance, but women who miscarried are somewhat less likely to have been born outside the Nordic countries. The largest difference lies in the age at labor market entry; women who experienced pregnancy loss were, on average, 0.74 years older when they entered the labor market. One possible explanation for this difference is that fecundity is declining with age, and women who enter the labor market at older ages are also somewhat older at the time of first pregnancy attempt. Thus, it is important to control for the age at labor market entrance. The independence assumption is thus made conditional on health and other observable characteristics.

Figure 2.1: Differences in average number of (pre-natal) hospital admissions by diagnosis, measured as the total number of hospital visits for each diagnosis category during the five years preceding first birth. The lower graph plots differences in average characteristics. Corresponding 95-percent intervals are also plotted.





Since hospitalizations reflect the most severe health issues, one might still worry that there are differences in health between the groups that are not captured by differences in hospitalizations. To get a less crude estimate of the differences in average health between the two groups, I therefore also examine the difference in health-related work absence. For this purpose, I use individual level data on sickness absence days from the Social Insurance Agency.[9] *Figure 2.2* plots the residuals from an Ordinary Least Squares (OLS) regression of the number of sickness absence days per year on year-fixed effects and age dummies. The two separate lines represent women who experienced pregnancy loss before first birth (solid line) and women who did not (dashed line). The x-axis displays the time since first birth for those who did not experience pregnancy loss, and time since miscarriage for those who did, i.e., the vertical zero-line approximates the year that they would have given birth to their first child, had they not miscarried. Importantly, the trends in sickness absence before first birth and before miscarriage are very similar for the two groups, with a small difference emerging a few years before birth/pregnancy loss. Moreover, the peak in sickness absence for those with no pregnancy loss reflect the increased absence associated with pregnancy. The sickness absence then decreases right after birth (which is the time period when they are on formal parental leave) to increase again and then stay rather constant.[10] For women who experienced pregnancy loss, there is an increase in sickness absence days in the year of the pregnancy loss, which likely is directly associated with the miscarriage. The high levels of sickness absence following this peak is most likely connected to childbirth, since this reflects the time period that this group actually have their (delayed) first birth. The reason for this increase seeming to be longer-run compared to women without pregnancy loss could simply be because the women in the former group have their first child at different times after pregnancy loss.

One might be worried here that the sickness absence associated with childbirth is

[9] From 1992 onwards, sickness absence spells are only recorded if they exceed 14 days; an employer-provided sick-pay was introduced in 1992 where the employer pays sickness benefits for the first 14 days of an absence spell, after which the benefits are paid from the Social Insurance Agency.

[10] This pattern of women's absenteeism by parenthood status is in line with the pattern recently documented by Angelov et al. (2013). Using Swedish data, the authors show that, before first birth, there are no differences in sickness absence between women and men. After entering parenthood, however, women increase their sickness absence by between 0.5 days per month more than their spouse, a difference which increases to 0.85 days more than the spouse in year 17 after first birth.

higher for women who miscarried compared to women who did not, since the peak in sickness absence is higher for the former. In this case, the IV estimates would be biased downwards. However, *Figure 2.2* also shows that the sickness absence of the two groups converge after 5-6 years after childbirth/miscarriage which tentatively suggests that miscarriages alone do not have any long-run impacts on health. In *Figure 2.3* I plot parameter estimates from an OLS regression of sickness absence days on miscarriage, by years since first birth or years since the first birth would have occurred had woman i not miscarried. Included controls are pre-motherhood number of hospitalizations, and hospitalizations by diagnosis type, an indicator for non-Nordic background and the age at labor market entrance as well as the calendar year of labor market entrance. The estimates confirm the findings from *Figure 2.2* and suggest that, after an initial period of higher sickness absence, the sickness absence of the two groups of women converge. This is in support of the exclusion restriction, where the concern is that mothers who experience a miscarriage might be adversely affected in that it induces psychological distress. Such a negative effect on health in turn might reduce a woman's hours worked, which would violate the exclusion restriction. Although the results presented in *Figure 2.2* do not indicate large long-run differences in sickness absence between the two groups of women, were this the case then 2SLS estimates would again be downward biased.

Lastly, an additional problem potentially inherent in using miscarriages as an instrument for birth timing was raised in a recent paper by Ashcraft et al. (2013) and may be important to keep in mind when interpreting the results in the present paper. They study the effect of teenage childbearing using miscarriages as a source of exogenous variation in birth timing. However, they argue that miscarriages are not socially random in the sense that willingness to abort reduces miscarriage risk. Moreover, teens who have abortions come from less disadvantaged backgrounds than those who do not. Thus, teens who miscarry are not a random sample of pregnant teenagers, but are from a more disadvantaged background. This implies that the IV estimator underestimates the the true costs of teenage childbearing. Thus, the authors conclude that when miscarriage is used as an instrument for birth timing, the estimates are biased towards a benign view of teenage childbearing.

Figure 2.2: The residuals from an OLS regression of sickness absence days (per year) on year-fixed effects and age, separately for mothers who experienced a miscarriage before first birth and who did not. The zero-line represents time since miscarriage for women who miscarried, and time since first birth for those who did not, respectively.

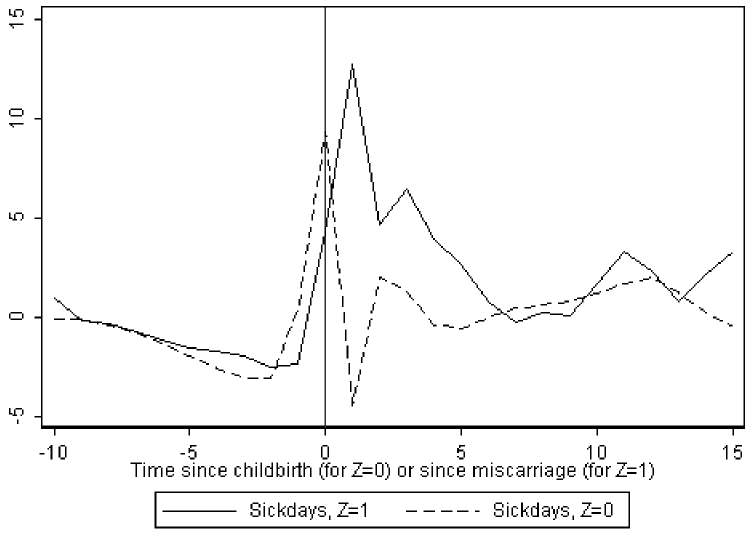
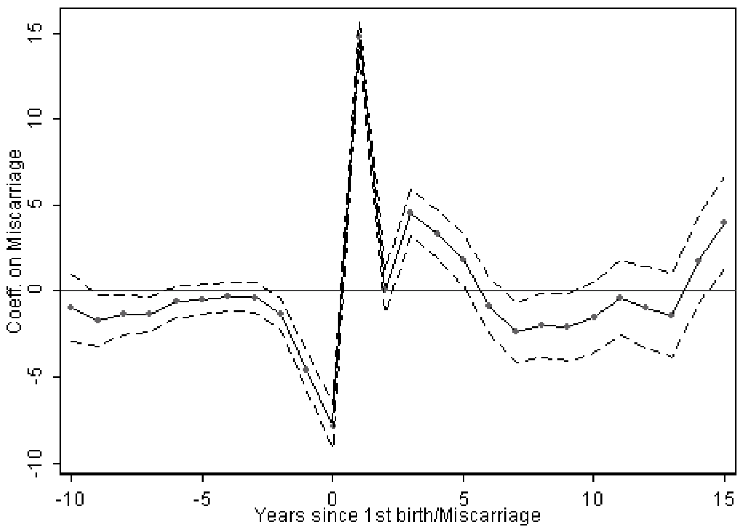


Figure 2.3: Parameter estimates from an OLS regression of sickness absence days (per year) on miscarriage before first birth. The x-axis represents time since first birth or time since the first birth would have occurred, had woman *i* not miscarried.



3 Data description and analysis sample

3.1 Data sources and definitions

The data used for the analysis is created by combining several Swedish population-wide registers. First, I use the multi-generation register, which links all children to their biological parents and provides information on birth year, birth month and birth order for each of all individuals' children. To these data I match registers containing information on a set of background characteristics such as age, gender, marital status, country of origin, highest attained educational level and graduation year, along with information on annual labor earnings from tax registers. Moreover, I add variables from a linked employer-employee data set providing information on the establishment at which the individual is employed each year, the first and last calendar month in a year that the worker receives income from the specific employer, information on industry affiliation, and the total income earned from the specific employer in that year. These registers cover the entire Swedish population aged 16-64 between 1985 and 2007. I then add individual level data on full-time equivalent monthly wages for each person-year-establishment observation, obtained from the Wage Structure Statistics and available for the entire public sector and about half of the private sector firms, for the time period 1985 through 2007. I also match individual level data on parental leave usage from the Social Insurance Agency for all individuals in my sample.

Finally, individual level data on miscarriages are provided by the National Patient Register (NPR). The NPR covers the universe of all hospitalizations (inpatient care) in Sweden. It includes medical information associated with each hospital visit, classified according to the International Classification Standard for Diseases (ICD). Using the ICD-codes, I can identify all hospital visits associated with miscarriages, for the time period 1987 through 2005.[11] Since the NPR does not record the order of the pregnancy for which the miscarriage occurred, I define the instrument - which indicates whether the *first* pregnancy ended in miscarriage - as being equal to unity if a miscarriage is recorded in the NPR for an individual before the birth year of her first child. Individuals with recurring

[11] The ICD-10 code for miscarriage is O.03.

miscarriages are entirely dropped from the sample.[12]

The inpatient record contains a non-negligible number of reported miscarriages. However, the number of reported cases decreases substantially each year from 1987 to 2005. *Figure A2* in the Appendix shows that the trend in reported miscarriages over time does follow the decreasing trend in the number of births in Sweden during the same time period. However, the number of miscarriages when illustrated as the fraction of births shows that the decrease is much larger than would be expected were it proportional to the decrease in the fertility rate. This is likely explained by technological change, i.e, changes in the medical treatments following a miscarriage and thereby in what type of medical establishment they are treated; recall that the NPR only includes inpatient care. Over time, it has become more common practice with medicinal treatment, as opposed to surgical treatment, following miscarriage, which do not have to be carried out in a hospital. Important to note is that this applies to to all health care in general. Thus, it is possible that surgical treatment is carried out as "day-surgeries" at outpatient care facilities. One potential concern is then that the cases of miscarriages that actually are treated at a hospital are more severe compared to the cases where treatment is acquired at an outpatient establishment. Women experiencing miscarriages with additional medical complications might be induced to reduce their working hours due to both medical and psychological reasons, which would violate the exclusion restriction. To explore this issue, I use study the frequency of reported co-morbidities for all hospital visits due to miscarriage. This is possible because the NPR reports not only main diagnosis for each hospital visit, but when relevant, up to 7 secondary diagnoses for co-morbidities. The frequency table *Table A1* in the Appendix shows that 95 percent of all miscarriages have no reported co-morbidities and 5 percent have 1 co-morbidity. There are very few cases in which more than one co-morbidity is reported. For individuals with a reported co-morbidity, I also tabulate the frequency by the medical causes for the first secondary diagnosis, shown in *Table A2*. As seen, the majority of the cases concern diseases of the genitourinary system or pregnancy-related diagnoses. However, even if there is not a high frequency of reported

[12] One might worry that recurring miscarriages induce psychological distress, which in turn might influence labor market outcomes directly, thereby violating the exclusion restriction.

co-morbidities to miscarriages, one could still worry that miscarriages that require hospital care are of greater medical severity compared to cases that do not show up in the inpatient records. As an additional analysis, I can also use the fact that the NPR provides a detailed description of the type of miscarriage, which I divide into four categories: complete, with and without complications and incomplete, with and without complications. This information on miscarriage type is, however, only available between 1997 and 2005. Nevertheless, using the available data I can attain an indication of the medical severity of the reported miscarriages. *Table A3* in the Appendix tabulates the occurrence of miscarriages divided into the four categories described above, among all miscarriages that are recorded in the inpatient records over the period 1997-2005. The results reported in *Table A3* show that the overwhelming majority - about 87 percent - are without complications. Moreover, *Figure A3* graphs the proportions of miscarriages with and without complications over the entire time period and shows that the majority of cases are without complications for all years, although the trends converge somewhat. Thus, although technological change has lead to fewer miscarriages being treated at inpatient establishments, there is no strong evidence towards only severe cases being treated at inpatient establishments as more and more cases are treated outside hospitals. Nevertheless, to control for technological change, I include year-fixed effects in all regressions. Furthermore, since I only have access to the inpatient records, an increasing number of "control" individuals will have had a miscarriage, but treated at an outpatient establishment, which implies that I will likely underestimate the effect of miscarriages on first birth timing, that is, the First-stage relationship is likely to be understated.

Since I define first-birth timing in terms of a woman's career - the time elapsed between labor market entry and the birth year of the first child - I need a clear definition of labor market entry. To identify each individual's first job and the associated starting wage, I define the "first" job as the employment that fulfills the following criteria: the first job (i) after completing the highest level of education which (ii) lasts for at least 4 months and (iii) yields annual earnings of at least 3 times the 10th percentile of the full wage distribution.[13] This definition of a first job is drawn upon the definition used in Kramarz &

[13]The latter is used to proxy a minimum wage as Sweden does not have a legislated minimum wage.

Skans (2014), although I use a different proxy for the minimum wage, and information on graduation year and educational attainment is here obtained from a different data source. *Figure A4* in the Appendix shows the time elapsed (cumulative) in order to find a first stable job for women with at most high school education and college education, respectively. For college educated women, about 60 percent find a job already in the same year as college completion and 80 percent find a job within one year of college graduation. For women with at most high school education, it takes significantly longer time to enter a stable employment: roughly 60 percent find a job within one year after high school graduation and about 80 percent find a first job within 3 years after completing high school.[14] Annual labor income and wages are all expressed in 2008 years prices (deflated using the Consumer Price Index).

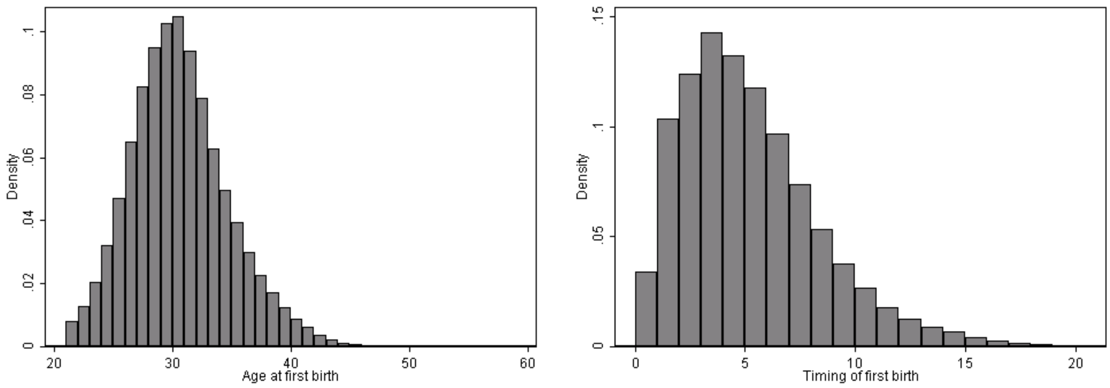
3.2 Analysis sample

I restrict attention to women who gave birth to their first child between 1988 and 2006. This population consists of 901,940 individuals, in total, of which 33,348 were reported to have experienced a miscarriage in the inpatient register some time during 1987 through 2005, with 20,207 of which the miscarriage happened before the birth of the *first* child. I restrict the sample to college educated women who were aged 21 or older at first birth and had their first child *after* entering the labor market. This restriction is made because, as noted by Herr (2011), the timing variable can only measure potential experience for women who have some pre-birth labor market experience. Furthermore, for those who experienced a miscarriage before first birth, I require that the miscarriage occurred after labor market entry, and I exclude those women who wait more than 10 years before having their first child, due to the monotonicity assumption not being satisfied for these women. In order to compare the restricted sample of highly educated mothers with the full sample of college educated mothers, *Table A4* also provides summary statistics for these two groups. As can be seen from *Table A4*, the sample restrictions leave me with a positively selected sample of highly educated women; compared to the full sample of college educated women, the women in the study sample are older when they have their

[14] These results are in line with findings presented in Kramarz & Skans (2014), from which the definition of a first stable job is drawn.

first child, they have more years of pre-birth labor market experience (4.3 years compared to 1.4 years), they are younger at labor market entry, find their first stable job sooner after completing college, are less likely to be married at labor market entry and less likely to be born outside the non-Nordic countries, and are more likely to live in a large city. Thus, the studied individuals might have stronger preferences towards market work than college educated in general. In *Figure 3.4* I plot the distribution of age at first birth and the dis-

Figure 3.4: Distribution of Age at first birth and Timing of first birth with respect to labor market entry, and years elapsed between pregnancy loss and first birth.



tribution of first birth timing in terms of the career for the analysis sample. As seen from *Figure 3.4*, the overwhelming majority of women had their first child within 10 years after entering the labor market.

Because I restrict the population of interest to women with at least one child, one question that arises is whether the occurrence of a miscarriage affects the probability to have a child at all, i.e., whether miscarriage affects the extensive margin of fertility. What would then be a cause for concern is whether individuals who miscarry and never become mothers differ from those who miscarry but subsequently give birth to a child. For the population of all Swedish women who were aged 45 or older in 2007 and who experienced a miscarriage between 1987 and 2005, *Table A5* in the Appendix reports summary statistics for women who had at least one child by the age of 45 and women who remained childless at the age of 45. First, we can note that there are a few statistically significant differences in average characteristics between mothers and childless women. For example, childless women have somewhat lower family incomes in 2007, likely attributed to the lower propensity to have been married. Moreover, childless women have somewhat lower own earnings, albeit not significantly different from mothers, but are weakly significantly more likely to have had a college education, and were on average older in 2007. While it is difficult to draw any clear conclusions regarding the potential selectiveness of the group of women that are excluded from the sample, i.e., women who had a miscarriage but remained childless, they seem to be a slightly negatively selected group in terms of own and family income. This would imply that the Reduced form estimates would be positively biased. However, among women who experienced a miscarriage, very few women - only four percent - remained childless by the age of 45. Hence, there is no immediate concern that conditioning the sample to include only mothers will bias the estimates in a significant way.

4 Results

4.1 The experience-wage and experience-income profiles of mothers

In this section I illustrate the labor income- and wage-experience profiles graphically for sub-samples of mothers with varying first birth timing. *Figure 4.5* plots the residuals from an OLS regression of annual earnings on year-fixed effects and dummies for age at labor market entry over the work history (where year 0 is the entry year in the labor market, defined as outlined in the Data section). Important to note is that labor earnings do not include parental leave benefits (or other transfers) and thus only measure income from market work. The five graphs in *Figure 4.5* represent five different groups of women defined by their timing of first birth, that is, by the number of years elapsed between their labor market entry and the birth of their first child: 0-1 years, 2-3, 4-5, 6-7 and 8-9 years. Evident from the figure is that, except for the group of women with first birth timing 0-1 years after entry, labor earnings for all groups of women are more or less identical in the year that they enter the labor market. Also, the income trajectories of the different groups follow each other remarkably closely until the first child is born, after which they diverge. This tentatively suggests that first birth timing may causally affect labor earnings. Secondly, 15 years after entering the labor market and beyond, women who gave birth to their first child as late as 8-9 years after labor market entrance have lower earnings than 'earlier' mothers. Thus, there is a permanently lower income for 'late' mothers after birth compared to earlier mothers. Furthermore, there is a striking downward shift in the income profile after birth for all groups, suggesting an almost permanent shift to part-time work following the birth of the first child. *Figure 4.6* graphs the evolution of the full-time equivalent monthly wage over the work history by first-birth timing, with the groups defined analogously to those presented in *Figure 4.5*. As was the case with labor earnings, also the wage profiles of mothers are strikingly similar until the first child is born, with the exception of mothers with very early childbearing (who have the lowest starting wage). Moreover, all groups seem to experience a slope decline in the wage path after they become mothers, which suggests reduced returns to experience post-birth, either due to reduced effort or due to reduced opportunities for on-the-job training or advancement. Interestingly, women with the lowest starting wages seem to be the ones postponing childbearing the longest; and they also seem to catch up in the long run with women who started at a higher wage level, but gave birth to their first child earlier.

Figure 4.5: The residuals from an OLS regression of annual labor income on year-fixed effects and age at labor market entry for five groups of women divided by their first-birth timing.

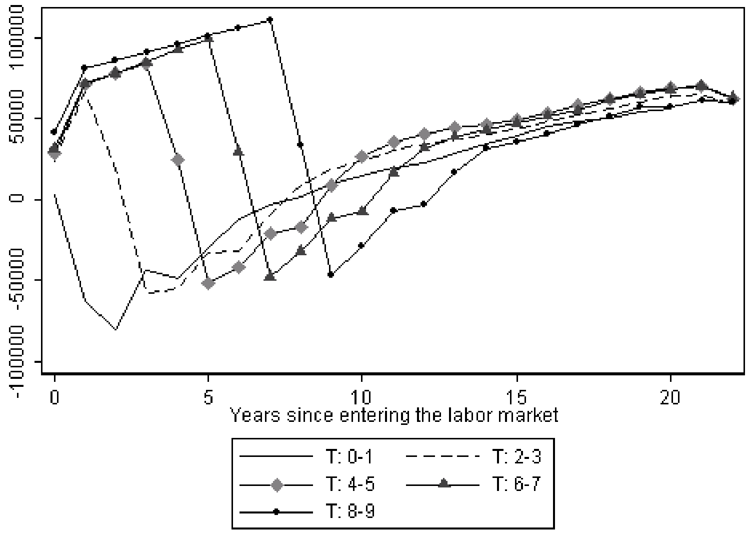
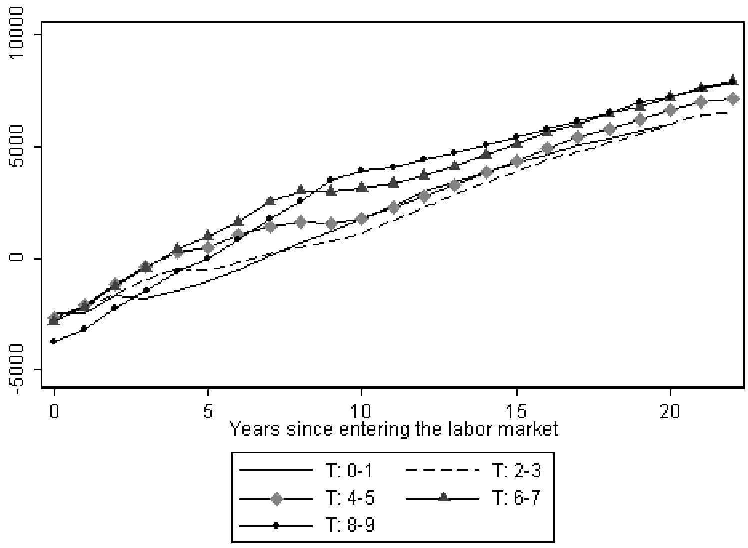


Figure 4.6: The residuals from an OLS regression of monthly wages on year-fixed effects and age at labor market entry for five groups of women divided by their first-birth timing.



4.2 The effect of motherhood delay on earnings and wages

Consider individual i, who entered the labor market in calendar year l, had her first child in calendar year b, and was a years old when she entered the labor market. The regressor of interest, T, is defined in terms of l as $T = b - l$, such that $T = 0, 1, ..., J$. Furthermore, the main outcome variable measures the natural log of total income earned over the first 20 years of individual i:s career. However, the data does not allow a 20-year follow-up period after labor market entrance for all individuals. Instead, career earnings are defined as the total income earned from the first year on the labor market up to as long as I can follow the individual, but at most up to 20 years. Thus, the estimated regression equation is:

$$ln(\sum_{l=0}^{L} y_{ial}) = \alpha_0 + \beta T_i + \delta_a a_i + \delta_L + \delta_l l_i + \mathbf{x}_i' \delta_x + \varepsilon_i \quad (1)$$

where a_i are dummies for age at labor market entry; δ_L are dummies for the number of observed career-years for individual i in the data, l_i are dummies for the calendar year of labor market entry (which are included to pick up e.g. wage growth in real terms), and finally, $\mathbf{x}_i$ is a vector of individual characteristics, measured pre-motherhood or at labor market entry. Equation (1) is estimated using both OLS and Two Stage Least Squares (2SLS) where T_i is instrumented with miscarriage before first birth. The coefficient of interest is β, which measures the average causal effect of a one-year delay of motherhood on the natural log of career earnings, or - in effect - the average impact of one extra year of pre-birth labor market experience. Note that a miscarriage induces a change in the birth timing for women, but also induces a change in the age at first birth. With the specification used here, I cannot identify an effect of postponing birth independent of the age at first birth. Although most predictions in the previous literature about the mechanisms of an effect of the age at first birth concern the level of pre-motherhood labor market involvement, I cannot rule out that also the age at first birth itself matters for outcomes. The effect measured here would then be a combined effect of pre-motherhood experience and the age at first birth.

Before analyzing the effect of first birth timing on income and wages, I first present evidence of the relevance of the instrument - miscarriage before first birth - for the first birth timing. *Table 4.1* depicts the OLS estimates of the effect of miscarriage on first birth timing, where the first column reports the results from a regression without covariates, and columns 2 to 4 present results from models where control variables are added stepwise. The results show an estimated delay of first birth timing by around 6 months in the model without control variables. Adding controls for age at labor market entry, non-Nordic background and marital status (column 2) decreases this estimate somewhat; pregnancy loss is then estimated to delay first birth by 5.1 months, on average. However, adding a control variable for the number of pre-natal hospitalizations (column 3) does not alter the estimate much, and neither does adding control variables for the number of pre-natal hospitalizations for different diagnoses (column 4). Column 5 shows the results from a regression where also a full set of dummy variables for the calendar year of labor market entry are included, as well as dummy variables for the number of observed career years. Including the calendar year dummies decreases the magnitude of the first-stage relationship quite considerably; the estimated effect of miscarriage on first birth timing now shows a delay of first birth of 2.4 months (0.2 years) including all relevant control variables. This is likely because an increasing number, over time, of women in the 'control' group are also experiencing a pregnancy loss, but are treated at an outpatient establishment. However, the F-statistic for joint significance in the first-stage (not shown) is 32.53, which is well above the suggested rule of thumb of 10. Thus, there is no concern of a weak instrument. In the following, I always include calendar year dummies in all regressions, as well as included in specification (5) of *Table 4.1*.

Table 4.1: OLS estimates of the effect of miscarriage before first birth on first-birth timing

	(1)	(2)	(3)	(4)	(5)
Miscarriage at first pregnancy	0.508***	0.424***	0.415***	0.393***	0.200***
	(0.037)	(0.036)	(0.036)	(0.036)	(0.035)
Non-Nordic background		-0.144***	-0.146***	-0.150***	0.164***
		(0.023)	(0.023)	(0.023)	(0.022)
Married at labor market entry		-2.274***	-2.277***	-2.279***	-1.764***
		(0.014)	(0.014)	(0.014)	(0.015)
Divorced at labor market entry		-1.180***	-1.193***	-1.205***	-0.475***
		(0.060)	(0.060)	(0.060)	(0.056)
No. of pre-natal hospitalizations			0.122***	-0.017	-0.012
			(0.014)	(0.015)	(0.013)
Additional controls					
Dummies for age at labor market entry		✓	✓	✓	✓
Hospitalizations by diagnosis				✓	✓
Dummies for calendar year of labor market entry					✓
Dummies for the no. of observed career-years					✓
Observations	223412	223412	223412	223412	223412

NOTES.—The outcome variable measures first-birth timing, defined as the number of years elapsed between labor market entry and first birth. Robust standard errors are reported in parentheses. *p<0.1, **p<0.05 ***p<0.01.

Table 4.2 depicts the results from the OLS and 2SLS estimations of the effect of first birth timing on the natural log of career earnings and the natural log of the average full-time equivalent monthly wage over the observed career, based on specification 1. As mentioned above, career earnings are defined as the total income earned from the first year on the labor market up to as long as I can follow the individuals in my data, at most up to 20 years after labor market entry. This outcome variable can be viewed as the net present value of income over the (observed) career. The average career wage is defined as the average full-time equivalent monthly wage (which is comparable to the hourly wage) over the observed career. Both income and wages are deflated using the Consumer Price Index.

Using OLS estimation, the results suggest that a one-year delay of motherhood is associated with an increase in career earnings with, on average, 3.7 percent. When instrumenting first birth timing with miscarriages, however, the 2SLS estimate suggests a statistically significant negative effect of a one-year delay on career earnings of about 15 percent. The standard errors of the 2SLS estimates are larger compared to the standard errors from the OLS regression, however, the F-statistic for joint significance in the first stage is 32.53, which is well above the suggested rule-of-thumb of 10. Thus, miscarriage before first birth does not seem to be a weak instrument. Moreover, the reduced form suggests a negative effect on earnings by 3 percent, significantly different from zero at the 1 percent level.

Labor earnings reflect both hours worked and hourly wage rates. To get a more complete picture of the career effects of birth timing, therefore, I continue by analyzing the effect of birth timing on the average monthly wage over the career, and the results from this analysis are presented in columns 3 and 4 of *Table 4.2*. Also here, the OLS estimate suggests a positive effect of delay, estimated to 1.8 percent higher wages, on average. However, instrumenting birth timing with miscarriages, the 2SLS estimate suggests a negative effect of a one-year delay on the average wage by 5.3 percent, significantly different from zero at the ten percent level. These results suggests that OLS exaggerates the positive effect of delay, and once endogeneity is taken into account, the effect of delay even goes in the opposite direction. This is in contrast to earlier studies who find positive effects of motherhood delay on both earnings and wages (see e.g. Miller 2011).

Up to this point the outcome variables measure income or wages over the entire, observed, labor market history for each individual. Hence, income both before and after entry into parenthood are included. As an alternative analysis I also consider the effects of postponing childbearing on the post-motherhood income trajectory, which also allows an examination of some dynamics of the effect of postponing childbearing. To this end, I perform separate yearly 2SLS regressions to estimate the effect of delayed childbearing on post-birth labor market income for varying years since first birth. The estimates from these regressions are plotted in *Figure 4.7*, which shows a large drop in income for mothers who postpone childbearing in years 0 to 3 after first birth. Income then 'bounces' back somewhat, but remains negative for the entire follow-up horizon (15 years after birth), although the estimates are then not significantly different from zero.

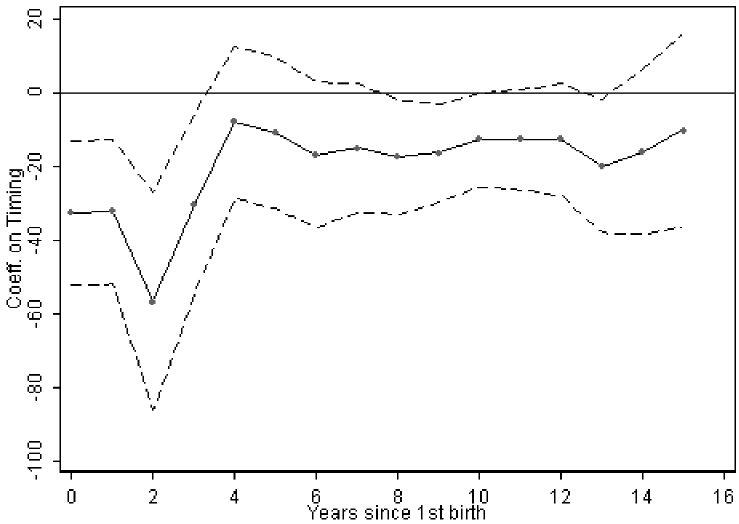
As seen from *Figure 4.7*, mothers who delay first birth due to the first pregnancy ending in miscarriage have a higher income drop also in the year of childbirth, which might cause some concern that these women would have lower income also before childbirth. However, the results presented in Section 2 did not indicate any major health differences between women who experienced pregnancy loss and those who did not. Nevertheless, as an additional sensitivity analysis, I perform separate regressions of the effect of pregnancy loss on labor income (i.e., the reduced form equation) for varying years since first birth, including pre-motherhood years. The estimates from these regressions are presented in *Figure A5* in the Appendix. As seen from this graph, while the effect of pregnancy loss on income is significantly negative for all years *after* birth, there are no large differences in income between women who miscarried and women who did not before motherhood. If anything, there is a tendency of a positive trend in income before motherhood for mothers who later experienced a pregnancy loss. This positive trend is, however, followed by a drop in income in the year before the birth of the first child. This is the year when most of the women in the sample who experienced pregnancy loss actually had the miscarriage, and could therefore be related to e.g. sickness absence associated with the pregnancy loss. It seems unlikely, however, that this income drop by itself drives the *long-run* negative effects of postponing motherhood on labor earnings. Rather, it seems like postponement of first birth causes a longer work interruption after birth. To try shed some light on what may cause the income drop after first birth, I therefore continue by investigating the effects of postponing motherhood on subsequent fertility behavior; the number of children and the spacing of subsequent births.

Table 4.2: OLS and 2SLS estimates of the effect of first-birth timing on the log of total earnings and the average wage rate over the first 20 years of the career

	Log Career earnings		Log Avg. career wages	
	OLS	IV	OLS	IV
Timing	0.037***	-0.147***	0.018***	-0.053*
	(0.000)	(0.050)	(0.000)	(0.029)
Non-Nordic background	-0.197***	-0.166***	-0.159***	-0.147***
	(0.007)	(0.011)	(0.005)	(0.007)
Married at labor market entry	-0.047***	-0.373***	-0.024***	-0.149***
	(0.004)	(0.088)	(0.003)	(0.051)
Divorced at labor market entry	-0.064***	-0.152***	-0.044***	-0.076***
	(0.017)	(0.031)	(0.012)	(0.018)
Pre-natal hospitalizations	-0.020***	-0.022***	-0.013***	-0.014***
	(0.004)	(0.005)	(0.003)	(0.003)
Dummies for calendar year of LM entry	✓	✓	✓	✓
Dummies for age at LM entry	✓	✓	✓	✓
Dummies for the no. of observed career-years	✓	✓	✓	✓
Number of hospital visits by diagnosis	✓	✓	✓	✓
Reduced form		-0.029***		-0.011**
		(0.008)		(0.006)
F-stat		32.5279		33.7912
Observations	223412	223412	222339	222339

NOTES.—The outcome variable measures the log of career earnings (columns 1 and 2), defined as the total income earned from labor market entry up to at most 20 years later, and the log of average career wages over the observed career (at most 20 years). Earnings and wages are deflated by Consumer Price Index. Robust standard errors are reported in parentheses. The control variables include dummies for non-Nordic born; married at labor market entry; divorced at labor market entry; number of hospitalizations pre-birth, number of hospital visits by diagnosis, as well as a full set of dummies for calendar year of labor market entry and the number of observed career-years. $^{*}p<0.1$, $^{**}p<0.05$ $^{***}p<0.01$.

Figure 4.7: Parameter estimates from separate 2SLS regressions of the effect of motherhood delay on labor income for varying years since first birth (where year 0 equals the birth year of the first child) and the 95 percent confidence intervals. The outcome variable measures annual labor income in 1000s SEK (expressed in real terms).



4.3 Exploring the mechanisms: subsequent fertility behavior

The results presented in the previous section suggest that a one-year delay of first birth causally induces a reduction of career earnings by 15 percent, on average, and a reduction in the average monthly wage over the career by 5 percent. In light of existing findings that often provide evidence of monetary benefits to postponing childbearing - in particular for highly educated women - these results may seem unexpected and surprising. In this section I aim to explore some of the potential mechanisms through which these effects on labor market outcomes could arise. The mechanisms proposed and analyzed here concern subsequent fertility behavior, both in terms of the total number of children and in terms of the spacing between the first and subsequent children. The former is interesting to analyze since it is possible that costs or benefits to postponing first birth partly capture a higher or lower wage penalty associated with more or fewer children, respectively. Moreover, in the demographics literature and in dynamic models of fertility, not only the timing of first birth is considered, but also the spacing of births, and how the timing of births relate to each other. For instance, in their paper on the timing and spacing of births using Swedish data, Heckman et al. (1985) find that, when controlling for unobserved heterogeneity, a delay in the arrival of one child is compensated for by an acceleration in the arrival of the next child. Furthermore, Troske & Voicu (2012) find that women with higher education have the first birth later in life, have fewer children, and space their subsequent children more closely together. Moreover, their findings suggest that spacing of births in turn affect women's labor market involvement.

To analyze whether completed fertility and child spacing are affected by delaying the first birth, I estimate the effect of motherhood delay on the total number of children born to a woman at the end of the observation period (i.e., in year 2007, which for many of the women in my sample represents completed fertility) and on the time interval between the first and the second child measured in years. The results from this analysis are presented in *Table 4.3*. The OLS estimation of the effect of motherhood delay on child spacing suggests that a one-year delay of the first birth reduces the spacing to the next child by roughly 2.3 months (0.19 years). Taking endogeneity into account, the 2SLS estimate suggests a reduction in the time interval between the first and the second child by about 8.4 months (0.70 years); an even larger effect compared to the OLS estimate. As the average interval between the first two births is about 2.7 years among women with more than one child in the sample, the reduction in the birth interval potentially implies relatively short birth intervals. However, when studying the effect on the total number of children born to a woman by the end of 2007, 2SLS estimation does not indicate that delay affects the total number of children. Closely related to subsequent childbearing is of course parental leave durations. Troske & Voicu (2012) conclude that while higher educated women have incentives to postpone second births, they are likely to space births more closely together. They further argue that this may suggest that higher educated women face larger fixed time or money costs of working, which makes them less likely to combine market work and child care. To study whether the first birth timing matters for the parental leave length of mothers, I estimate the effect of postponing first birth on yearly parental leave usage, starting from the birth year of the first child up to 7 years later and on annual fertility for the same years. Annual fertility is here defined as a dummy variable taking the value one when a second or third child is born. The results from this analysis are presented in

Table 4.3: OLS and 2SLS estimates of the effect of first-birth timing on child spacing and number of children

	OLS	2SLS
Dependent variable		
Years between 1st and 2nd child	-0.19***	-0.70***
	(0.00)	(0.16)
Number of children in 2007	-0.10***	0.02
	(0.00)	(0.06)
Controls for personal characteristics	✓	✓
Dummies for calendar year of LM entry	✓	✓
Dummies for age at LM entry	✓	✓
Dummies for the no. of observed career-years	✓	✓
Observations	223412	223412

NOTES.—The outcome variables measure the number of years between the first and second child, and the total number of children to a woman at the end of the observation period (2007), respectively. Each quadrant in the table thus presents the results from separate regressions. The control variables include dummies for non-Nordic born; married at labor market entry; divorced at labor market entry; and number of hospitalizations pre-birth, as well as a full set of dummies for calendar year of labor market entry and the number of observed career-years. *p<0.1, **p<0.05 ***p<0.01.

Table 4.4 and show an interesting pattern. First, we can see that in the first four years after birth, mothers who delay have on average 28, 34, 131, and 77 days more parental leave, respectively. Interestingly, they have 90 days less parental leave days, on average, 4 years after the first child is born. The coefficients for the remaining years are also negative, but not significantly different from zero. Finally, in the lower panel of *Table 4.4* I present the result from estimating a 2SLS regression of the effect of motherhood delay on the total number of parental leave days taken during the first eight years after the birth of the first child. This effect is estimated to 55 days, but is not statistically significant. Thus, it seems like postponing first birth leads to a reshuffling of parental leave, rather than maybe a total increase in leave taking. Consistent with this pattern are the results from the estimates of the effect of delay on annual fertility (for subsequent children) which are presented in column 2 of *Table 4.4*, where we see an increase in fertility of 14 and 56 percentage points 1 and 2 years after the birth year of the first child, respectively, and then a decrease of 32 and 13 percentage points 3 and 4 years after first birth, respectively, with remaining coefficients also being negative but not significantly different from zero. The time pattern of the effects on subsequent fertility are thus in line with the time pattern of the effect of postponing first birth on labor income shown in *Figure 4.7*.

The results on the spacing of births and on the total number of children are in line with some previous findings of the effect of motherhood delay on subsequent childbearing. For example, Bratti & Tatsiramos (2012) study the effect of delaying motherhood on the transition to the second birth for a number of European countries using data from the European Community Household Panel. The effect of delaying motherhood is found to differ across countries. Specifically, women who delay their first birth are less likely to progress to second parity, but a higher availability of family friendly policies raises the probability of having a second birth. For instance, the authors find that delaying age at motherhood from 25 to 30 leads to a positive effect on the likelihood of progressing to higher parity within 5 years from first birth in countries such as Denmark, and a negative effect of 12 percentage points in Southern European countries such as Greece.

Taken together, one possible interpretation of the results presented in this section is that delaying motherhood does not lead to fewer children altogether, but more closely spaced children, and that this, in turn, implies being away from the workplace during a larger part of one section of the working history, perhaps the critical time period of career build-up.

Table 4.4: 2SLS estimates of the effects of first birth timing on yearly parental leave usage and yearly probabilities of having another child

Dependent variable	Parental leave days	Additional child
Specification	2SLS	2SLS
Birth year first child	27.884*	0.002
	(16.416)	(0.021)
Birth year first child +1	33.396**	0.142***
	(16.609)	(0.051)
Birth year first child +2	130.939***	0.559***
	(40.989)	(0.165)
Birth year first child +3	77.454**	-0.322***
	(37.520)	(0.112)
Birth year first child +4	-90.274***	-0.126*
	(33.930)	(0.073)
Birth year first child +5	-31.689	-0.019
	(23.099)	(0.051)
Birth year first child +6	-19.399	-0.006
	(17.528)	(0.041)
Birth year first child +7	-2.917	0.009
	(13.688)	(0.029)
Birth year first child +8	77.064	0.013
	(65.890)	(0.051)
Pooled data		
Year 1 to year 8	54.965	
	(58.967)	

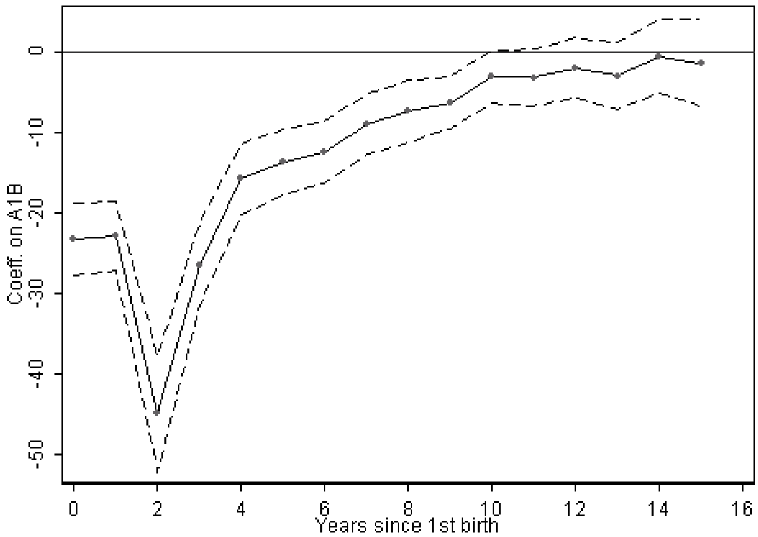
NOTES.— The outcome variables measure the number of parental leave days taken in each year from the birth-year of the first child up to 7 years later (column 1) and the annual probability of having an additional child, respectively. The control variables include dummies for non-Nordic born; married at labor market entry; divorced at labor market entry; and number of hospitalizations pre-birth, as well as a full set of dummies for calendar year of labor market entry and the number of observed career-years. Robust standard errors are reported in parentheses. *p<0.1, **p<0.05 ***p<0.01.

4.4 Robustness analysis: estimates of the effect of the age at first birth

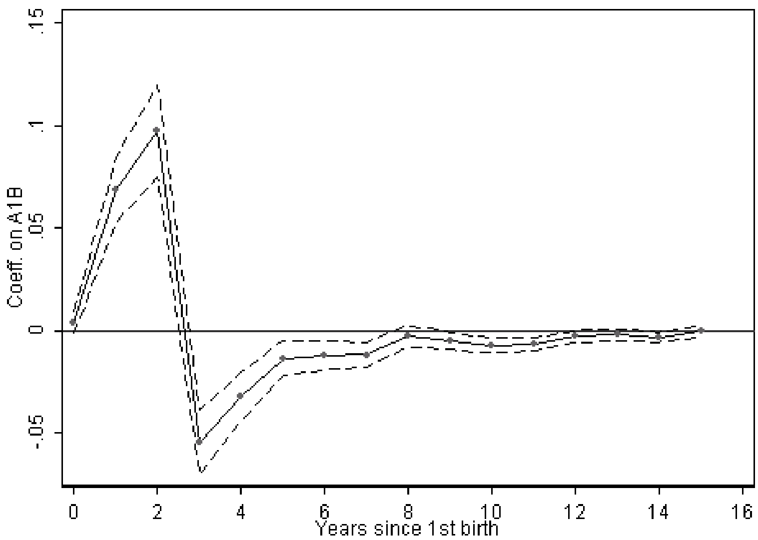
The estimates of the effect of delayed motherhood on labor market outcomes provided in this paper contrasts the results from previous work on the topic. While an analysis of the effect of delayed motherhood on subsequent fertility tentatively suggests that the tempo of subsequent births may partly explain the findings presented in this paper, there still might remain some concerns, for example about whether the results are driven by the somewhat different sampling scheme and the different definition of first birth timing employed. In this section, I therefore present results from analyses where I follow a more traditional sampling scheme. I now restrict attention to mothers whose first child was born between 1988 and 2006, who were aged 21 or older at first birth, but not older than 35 at first birth. Individuals with at most compulsory schooling are excluded from the sample. Thus, in contrast to my main analysis, I do not put any restrictions on the first birth being born before or after labor market entrance. The regressor of interest is here defined as the age at first birth, which is the most commonly used variable to measure first birth timing in existing work. I then estimate the effect of the age at first birth, using miscarriage before first birth to instrument for the former, on yearly labor income over a 15-year horizon after first birth. I also estimate the effect of the age at first birth on the annual probability to give birth to a subsequent child. The control variables include an indicator for non-Nordic background, the number of pre-motherhood hospitalizations, the number of pre-motherhood hospitalizations for each diagnosis type, and an indicator for college education. To conserve space, I present the findings graphically, but the full results are available upon request. As seen from *Figure 4.8*, the results look very similar to my main sample of interest. Specifically, there is a negative effect of postponing motherhood by one year, on average, with a sharp drop in earnings in the second year after first birth. This drop in earnings coincides with a peak in the probability to give birth to a subsequent child. All in all, measuring timing of birth in terms of the career point instead of as the age at first birth does not seem to be what is driving the negative effects of motherhood delay on earnings.

An additional concern regarding the negative effects found on earnings from postponing motherhood, is that the exclusion restriction is not satisfied. While there does not seem to be any long-run effects of miscarriage on sickness absenteeism, an increase in sickness absence appears in the year of the miscarriage event. In the next and final section of the paper, I therefore estimate the effect of first birth timing under a different set of identifying assumptions. Specifically, I estimate the effect of motherhood itself on wages, and allow this effect to vary by birth timing. This is achieved with panel data and an individual-fixed effects estimator.

Figure 4.8: The effect of Age at first birth on labor earnings, and on higher order fertility, mothers with at most high school or college education.



(a) 2SLS estimates of the effect of age at first birth on labor earnings, 1000s SEK



(b) 2SLS estimates of the effect of age at first birth on the annual probability of giving birth to a subsequent child

5 The effect of spacing births on earnings

The results so far suggest that postponing entry into motherhood negatively affects women's subsequent earnings. Motherhood delay was also found to accelerate the time to the next birth. I hypothesized, therefore, that the shorter spacing of births could potentially be the driving mechanism for the negative effects of motherhood delay. Spacing births may alleviate adverse consequences of further childbearing through the accumulation of pre-birth labor market experience. In addition, by allowing re-entry into the labor market between births, spacing births could imply work interruptions of shorter durations. In this section, I explore this mechanism further, by estimating the effect of birth spacing on subsequent income and wages. To isolate the causal impact of birth spacing, I exploit variation in spacing induced by miscarriages between the first and second live births. Using the same data sources as above, I estimate the effects for college educated women with at least two children, who were older than 21 at first birth, and who entered the labor market before first birth (thus the same data restrictions as the main analysis, but now conditioning the sample to include women with two children or more). I further restrict the sample to women who did not experience a miscarriage before first birth. Summary statistics for the study sample are reported in *Table A6* in the Appendix and show that the sample is similar to the group studied in the above analyses. For instance, 2.4 percent of the sample experienced a miscarriage between the first two live births, whereas 1.9 percent of the main sample experienced a miscarriage before first birth. Furthermore, the women in this study sample were on average aged 29.32 years at first birth, and the average spacing between the first two children - measured as the number of years elapsed between first and second birth - is 2.8 years.

Table 5.5 reports the results from an OLS estimation of the first-stage relationship as well as OLS and 2SLS estimates of the effects of birth spacing on the income and log wage rates for varying years after second birth. The coefficient on the instrument, presented in panel A of *Table 5.5*, suggests that a miscarriage between the first and second live birth increases the birth spacing between the first two children by roughly 0.77 years (9.2 months), on average. Thus, the first-stage effect is non-negligible. Panel B reports the results from the OLS and 2SLS estimation of the effect of child spacing on subsequent labor income. The table reports estimates for 2, 4, 6, 8 and 10 years after second birth, and the results indicate that postponing second birth by one year, on average, increases labor income, both when estimated in OLS and 2SLS. Furthermore, the positive effect is almost monotonously increasing with time since second birth. Panel C presents the results for wage rates, and show that also wages are positively affected, with the effect ranging between 3 and 4 percent over the first 2 to 10 years after second birth.

An assessment of the validity of the independence assumption is performed by estimating the reduced form equation on labor income (expressed in 1000s SEK) in years prior to first birth. The results from this falsification test are presented in *Table A7* in the Appendix and show that there are no differences in labor earnings before becoming parents between women who later had a miscarriage and women who did not; the estimated coefficients are small in magnitude and not statistically significant.[15] The effect of spacing births on labor income is thus found to be positive, sizeable and increasing in magnitude by time since birth. This is in line with the hypothesis that the shorter birth intervals caused by a delay in motherhood entry are a likely channel for the negative effects of postponed motherhood on earnings. One possible explanation for this finding is that postponing second birth induces mothers to re-enter the labor market between births to a greater extent, thereby gaining more labor market experience before the birth of the second child. Thus, spacing births in a longer interval could potentially imply a shorter consecutive absence from work for child care reasons, and a stronger labor market attachment as a result. It should be noted that the Swedish parental leave system allows relatively lengthy withdrawals from the labor market with wage replacement. In addition, the "speed premium" rules imply that individuals who have a birth interval of less than 30 months do not have to re-enter the labor market between births to re-establish eligibility to paid leave. It is therefore possible for parents, in particular mother, to be absent from the labor market for an extended time if two children are born in close connection.

[15] Importantly, in contrast to the results from the analysis of the timing of first birth, the reduced form estimates for this analysis on post second birth income are positive. These results are available upon request.

Table 5.5: First-stage relationship and the effects of birth spacing on income and log wage rates

	OLS	2SLS	N
A. First-stage relationship			
Miscarriage	0.767***	-	208800
	(0.024)	-	
B. Estimates of the impact of spacing on income, 1000s SEK			
Birth year +2	9.536***	4.759*	170651
	(0.233)	(2.447)	
Birth year +4	11.680***	7.135**	142471
	(0.289)	(2.840)	
Birth year +6	13.090***	13.845***	119170
	(0.386)	(3.398)	
Birth year +8	12.361***	11.600***	100441
	(0.444)	(3.772)	
Birth year +10	12.553***	16.320***	83438
	(0.541)	(4.323)	
C. Estimates of the impact of spacing on log wage rates			
Birth year +2	0.032***	0.030***	149169
	(0.000)	(0.004)	
Birth year +4	0.033***	0.036***	126194
	(0.000)	(0.005)	
Birth year +6	0.034***	0.037***	106050
	(0.001)	(0.005)	
Birth year +8	0.033***	0.040***	89307
	(0.001)	(0.006)	
Birth year +10	0.030***	0.035***	74153
	(0.001)	(0.006)	

NOTES.— The outcome variables measure labor earnings in 1000's SEK and the natural log of full-time equivalent monthly wage rates at 2, 4, 6 and 8 years after second birth, respectively. Labor earnings are deflated with CPI (and expressed in 2008 prices). Included covariates are the number of pre-natal hospitalizations, a dummy for non-Nordic background, dummies for high school education and college education, a full set of dummies for age at first birth and dummies for cohort. Robust standard errors are reported in parentheses. *p<0.1, **p<0.05 ***p<0.01.

6 Panel data estimates of the effect of motherhood on wages

The previous section presented results from estimating the effects of motherhood timing on the net present value of income over the observed career and on the average monthly wage over the career, using cross-sectional variation. These labor market outcomes were both found to be negatively affected by a one-year delay of motherhood. Moreover, I found that motherhood delay also induces a closer spacing between the first and the second born children, whereas no effects were found on the total number of children born by the end of the observation period. This might explain the negative effects on labor market outcomes, if being away from the labor market for a larger share of a critical time period is associated with larger wage penalties. Indeed, estimates of the effect of birth spacing on subsequent earnings support the conjecture that a short birth interval may have adverse consequences for women's career outcomes. In this section, I aim to explore the effects of motherhood on individual wage growth more closely, and how it varies with first birth timing. To this end, I estimate the effects of motherhood on full-time equivalent monthly wages using panel data specifications in the spirit of Miller (2011). In the first panel specification I include individual-fixed effects, experience (years worked), experience squared, motherhood status (set to equal one in the year of first childbirth onwards) and years since first birth. The motherhood indicator captures a potential downward shift of the wage profile, which Miller (2011) refers to as human capital depreciation or fixed motherhood penalties, whereas a negative coefficient on Year Since First Birth captures a slope decline in the wage profile post-birth, indicating reduced returns to experience. The estimated panel data specification is thus the following:

$$ln(w_{it}) = \beta_0 + \beta_1 Exp_{it} + \beta_2 ExpSq_{it} + \beta_3 Mother_{it}$$

$$+ \beta_4 Mother_{it} \times (YearsSinceFirstBirth_{it}) + \alpha_i + \varepsilon_{it} \qquad (2)$$

where Years Since First Birth is measured as the calendar year minus the calendar year of the first birth. α_i capture unobserved time-invariant individual-fixed effects and thus controls for unobserved individual heterogeneity. In a second panel data specification, the effect of motherhood is allowed to vary with first birth timing, by including interaction terms between years since first birth, motherhood and dummy variables for four groups of women with different birth timing:

$$ln(w_{it}) = \beta_0 + \beta_1 Exp_{it} + \beta_2 ExpSq_{it} + \beta_3 Mother_{it} + \sum_{j=1}^{4} \beta_{4,j} Mother_{it}$$

$$\times (YearsSinceFirstBirth_{it}) \times 1(Timing_i = j) + \alpha_i + \varepsilon_{it} \tag{3}$$

where $j = 0-2, 3-4, 5-6, 7-10$.

Table 6.6 reports the results from estimating Equations (2) and (3) in columns 1 and 2, respectively. As seen from *Table 6.6*, the results from estimating specification (2) indicate that wages increase with experience, by on average 5 percent per extra year worked. In addition, there is a fixed motherhood wage penalty of about 3.8 percent (i.e., a shift of the profile, represented by the coefficient on the motherhood indicator). Moreover, the interaction term between motherhood and years since first birth is negative and statistically significant with a coefficient of -0.0107, suggesting that mothers indeed experience a flattening of the wage profile post birth. When the changes in wages for mothers are allowed to vary with first birth timing, we see that the reduced return to experience is larger for women who delay motherhood for longer times; the coefficients are always negative and almost monotonously more negative for each group of 'delay'.

Table 6.6: Panel estimates of the effect of motherhood on log wages

	(1) FE	(2) FE
$PotentialExperience$	0.0502***	0.0498***
	(0.0002)	(0.0002)
$ExperienceSquared$	-0.0003***	-0.0002***
	(0.0000)	(0.0000)
$Mother$	-0.0376***	-0.0335***
	(0.0005)	(0.0005)
$Mother \times$ (Years Since birth)	-0.0107***	
	(0.0002)	
$Mother \times$ (Years Since birth) $\times 1(T = 0-2)$		-0.0108***
		(0.0002)
$Mother \times$ (Years Since birth) $\times 1(T = 3-4)$		-0.0130***
		(0.0003)
$Mother \times$ (Years Since birth) $\times 1(T = 5-6)$		-0.0124***
		(0.0003)
$Mother \times$ (Years Since birth) $\times 1(T = 7-10)$		-0.0141***
		(0.0003)
$Constant$	9.7055***	9.7050***
	(0.0005)	(0.0005)
Observations	1724014	1724014

NOTES.— Standard errors are clustered at the individual level and reported in parentheses. *p<0.1, **p<0.05 ***p<0.01.

Consistent with the main results presented in the previous section, these findings suggest that the wage penalty is larger for women who have their children later. The analysis based on the panel data set corroborates the negative effects from the analyses based on the cross-sectional variation, and indicate the presence of reduced returns to experience post-birth for mothers, with this slope decline in wages being increasingly larger for “late” mothers. This decreased returns to experience may either be attributed to fewer opportunities for advancement and training (the so called “mommy-track”), or women may exert less effort in the workplace.

7 Concluding discussion

The negative effects of career interruptions due to childbearing on women's employment and wages are rather well documented. However, the implications of first birth timing on career outcomes are not yet fully understood. In this paper, I aim to estimate the causal effect of postponing first birth on the labor earnings and wages of Swedish college educated women. To isolate the impact of first birth timing on labor market outcomes, I instrument fertility timing with the occurrence of miscarriage before first birth. I focus on highly educated women who first finish college, enter the labor market and subsequently become a parent. This allows me to measure birth timing as potential pre-natal labor market experience, i.e,. the number of years elapsed between labor market entry and the birth of the first child.

In line with previous studies, I find that OLS estimation suggest a positive effect of postponing first birth on the total income earned over the first 20 years of the career, as well as on the average monthly wage over the career. However, exploiting the arguably exogenous variation in birth timing induced by pregnancy loss, I find that postponing first birth by one year on average negatively affects both career earnings and wages. This is in stark contrast to most of the previous studies, who often document monetary benefits to postponing motherhood. I also find that delaying first birth causally reduces the time elapsed to second birth, that is, a decreased spacing between the first two children. However, I find no evidence of an effect of first birth timing on the total number of children. Thus, mothers who delay first birth do not seem to forego further childbearing, but rather to have children more tightly spaced. Closely linked to the findings on subsequent fertility, postponing first birth induces a 'reshuffling' of parental leave usage to be higher in the years closest to the first born child, and lower in the following years. Hence, fertility delay seem to induce mothers to have two (or more) lengthy parental leave periods more closely spaced, which in turn might be more detrimental to subsequent career opportunities compared to taking the same amount of leave but spread out over a longer horizon of working life.

To further study whether the shorter birth spacing may indeed explain the negative effects of motherhood postponement on earnings, I exploit the instrument to explicitly analyze the impact of birth spacing on subsequent earnings. To this end, I use miscarriage between the first and second live birth as an instrument for birth spacing, and estimate the effect of (longer) birth spacing on income and wages. The results from this analysis are in line with the hypothesis that short birth intervals have detrimental effects on subsequent career outcomes, as spacing births are found to have sizeable and positive effects on both income and wage rates lasting even 15 years after the second child is born.

To study how the individual wage pattern is affected by motherhood and the timing of motherhood, I also estimate the effect of motherhood itself on wage growth for women, using a panel data specification with individual-fixed effects. The panel data results corroborate the negative effects from the cross-sectional data set and suggest a slope decline of the wage profile post-birth, with the slope decline being increasingly larger for late mothers.

The findings provided in this paper are in contrast to studies from the US where fertility delay has generally been found to have positive effects on both income and wages. However, it is important to interpret the results in the present study within the institutional context. In Sweden, all parents are entitled to 480 days of job-protected parental leave for each child and in practice, most women take out a major part of this leave. The benefits are wage-replaced and parents have the right to reduce working hours for up to 25 percent until the child turns eight years old. Hence, Swedish parents do not face the same restrictions to take parental leave as do parents in e.g. the US. In a system that does not offer very generous family policies, establishing a stronger attachment to the labor market before taking leave to care for a child may protect mothers from having to start over when they re-enter the labor market, and also facilitate the possibility of returning to the pre-birth employer. Delaying motherhood in such a system may also partly reflect a lower wage penalty associated with fewer children. In Sweden, however, job protection is the default, and legislated parental leave is rather generous. Hence, long interruptions, especially if they are closely spaced, might yield stronger penalties in the labor market, at least for college educated women who are the focus of this paper. This of course raises questions about the optimal length of leave, and the potential detrimental effects for mothers' careers of having a too generous parental leave system, in particular if women continue to stand for the majority of parental leave take-up. The findings presented in the present paper and in the emerging literature about so called tempo effects of fertility thus may have important policy implications since, not only policies affecting the number of children, but also tempo policies that affects the age at first birth and spacing of births may have unanticipated implications on women's wage trajectories, and should provide interesting avenues for future research.

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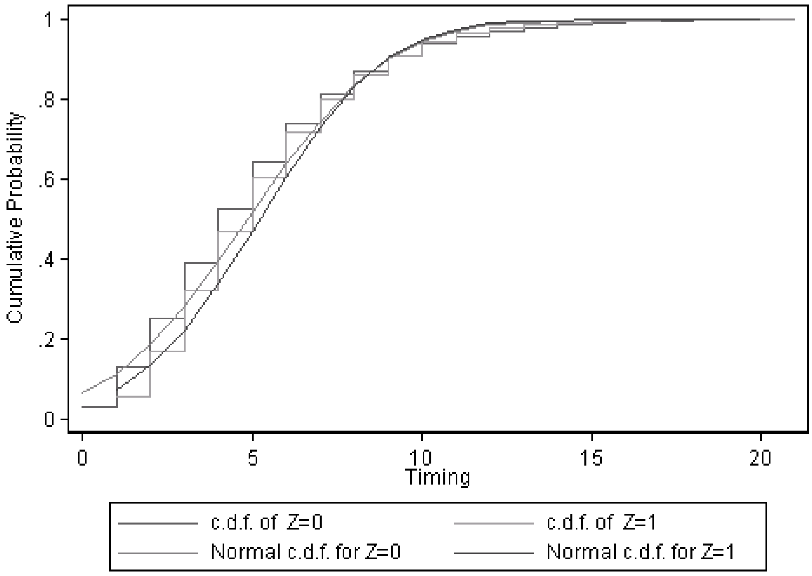
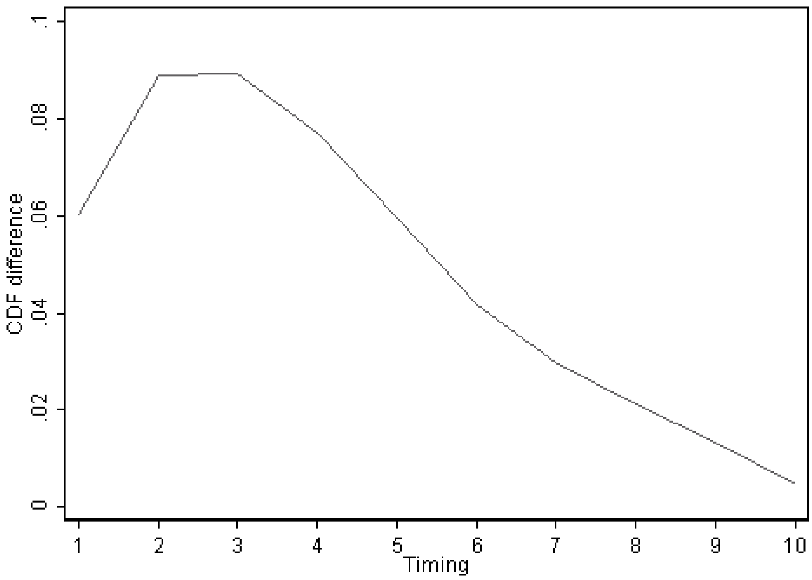
Appendix

Figure A1: The empirical Cumulative Distribution Functions of birth timing, T, by the occurrence of miscarriage before first birth.



(a) Empirical CDF of first birth timing for $Z = 1$ and $Z = 0$ and the best fitting normal model superimposed over the sample CDFs.



(b) Difference between the CDFs graphed in (a)

Figure A2: Number of reported miscarriages in the NPR, number of children born, and the share of miscarriages of children born, 1987-2005.

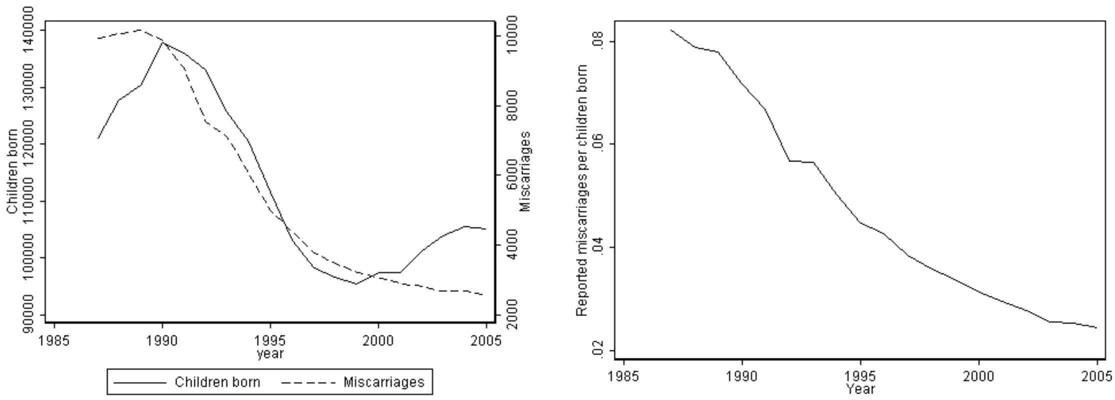


Figure A3: Proportion of miscarriages with and without complications, 1997-2005.

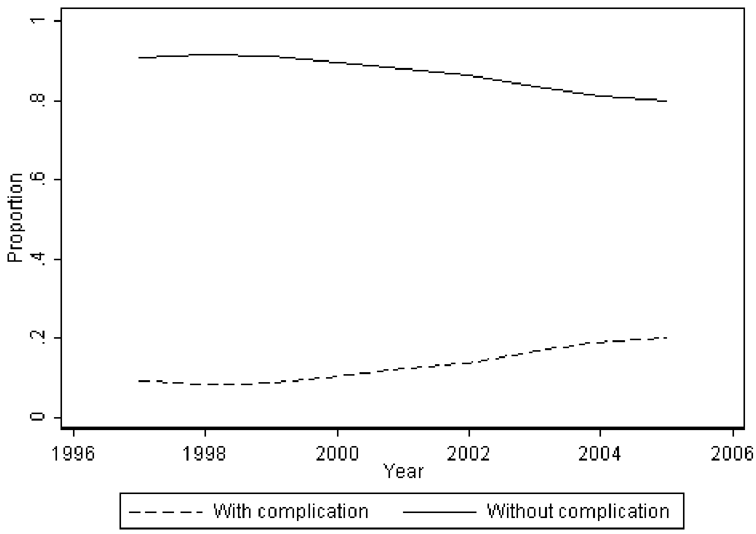


Table A1: The frequency of co-morbidities for hospital visits associated with miscarriage

	Frequency	Percent
No. of co-morbidities		
0	80,209	95.27
1	3,554	4.22
2	356	0.42
3	63	0.07
4	6	0.01
5	2	0.00
6	1	0.00

NOTES.— The table reports the frequency of co-morbidities to all hospitalizations where the main diagnosis is classified as a miscarriage.

Table A2: Diagnosis type for co-morbidities of miscarriage

	Frequency	Percent
Co-morbidity type		
Infectious	123	3.46
Tumors and neoplasms	255	7.18
Blood(-forming organs)	468	13.17
Endocrine	148	4.16
Mental behavioral	36	1.01
Nervous system	29	0.82
Ear	3	0.08
Circulatory system	40	1.13
Respiratory	58	1.63
Digestive system	28	0.79
Skin	15	0.42
Musculoskeletal	34	0.96
Genitourinary system	556	15.64
Pregnancy related	775	21.81
Perinatal	3	0.08
Congenital malformations	59	1.66
Symptoms not classified elsewhere	87	2.45
Factors associated with health status	775	21.81
External causes	62	1.74

NOTES.— The table reports the frequency of the diagnoses of the first co-morbidity (i.e., the first secondary diagnosis to the main diagnosis being miscarriage) to all reported miscarriages with at least one reported co-morbidity.

Table *Table A3* divides the reported miscarriages into four different types based on severity. The data is based on reported miscarriages that occurred during 1997 to 2005. The table shows that the majority of cases are without any complications (adding both complete and incomplete miscarriages). Moreover, Figure *Figure A3* graphs these proportions, now divided only into two categories: with and without complications, by year. We can

Table A3: Severity of miscarriages, 1997-2005

	Mean
Incomplete with complication	0.108
	(0.311)
Complete with complication	0.0183
	(0.134)
Incomplete without complication	0.676
	(0.468)
Complete without complication	0.197
	(0.398)
Observations	26120

NOTES.— Means and (standard deviations).

see that the proportions of reported cases with and without complications seem to converge somewhat over the time period, but the overwhelming majority of miscarriages are reported to have been without any complications throughout the time period.

Defining labor market entry

For individuals with more than one employment at the same firm in the same year I calculate the total earnings received from that employer in that year, as well as the total number of months worked and then drop duplicate observations on person-firm-year. Moreover, for individuals with more than one employment in one year, but at different firms, I define that individual's workplace the firm at which she received her main income in that year. This procedure leaves me with unique observations on person-firm-year level.

Since timing of motherhood is here defined as timing with respect to labor market entry, a definition of a first stable employment is needed. To do this, I first back out the year in which the highest attained education level is completed for each individual, where I divide highest attained educational attainment into three categories: compulsory education; high school education; and some college or more. Backing out the graduation year from the panel data gives an average age at completion of compulsory education of about 16.14, which is in line with Swedish compulsory education being 9 years of duration starting at the age of 7. For high school graduates, the corresponding age in the data is about 19.90 and for college educated 27.55 (the high average age for finishing college is partly explained by gap years between high school and college). Second, I define entry to the labor market as the first calendar year after the completion of education that the individual (i) earned at least three times the 10th percentile of the full wage distribution, and (ii) had an employment that lasted at least 4 months. Figure *Figure A4* shows the time elapsed between finishing education until a first job is attained for high school educated mothers and college educated mothers, respectively.

Figure A4: Time elapsed from graduation to a first job for high school and college educated mothers.

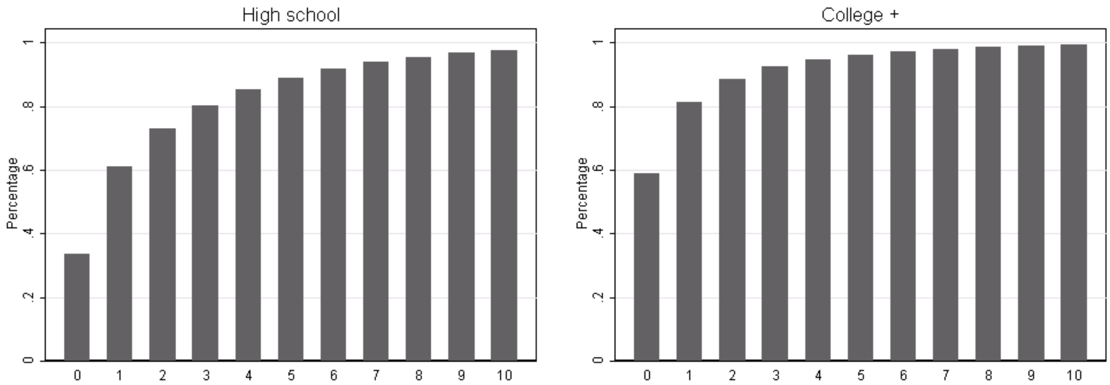


Figure A5: Parameter estimates from separate OLS regressions of the effect of pregnancy loss on labor income (the reduced form equations) for varying years since first birth (where year 0 equals the birth year of the first child) and the 95 percent confidence intervals. The outcome variable measures annual labor income in 1000s SEK (expressed in real terms).

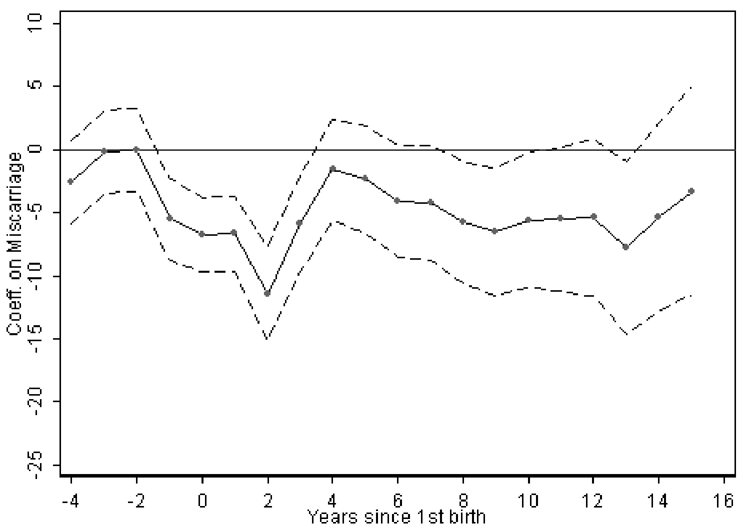


Table A4: Summary statistics for the study-sample, the sample of college educated women, and the full sample of women who had their first child 1988-2006

	Study sample	All college graduates	Full sample
Miscarriage at first pregnancy	0.0185	0.0208	0.0224
	(0.135)	(0.143)	(0.148)
Age at miscarriage	30.73	29.02	27.18
	(4.408)	(5.070)	(5.268)
Age at first birth	30.13	29.19	27.48
	(3.664)	(4.504)	(4.985)
1st Birth Timing	4.291	1.392	3.095
	(2.507)	(6.612)	(6.193)
Age at labor market entry	25.84	28.03	24.74
	(3.363)	(5.483)	(5.727)
Time to labor market entry	0.896	1.072	1.680
	(1.696)	(2.230)	(2.739)
Non-Nordic background	0.0486	0.139	0.164
	(0.215)	(0.346)	(0.371)
Married at labor market entry	0.0933	0.187	0.116
	(0.291)	(0.390)	(0.320)
Live in large city	0.293	0.265	0.209
	(0.455)	(0.441)	(0.406)
Number of children in 2007	1.955	1.957	1.990
	(0.710)	(0.763)	(0.842)
Compulsory schooling			0.0901
			(0.286)
High school			0.472
			(0.499)
College			0.438
			(0.496)
Observations	223412	382439	901940

NOTES.— The table reports summary statistics for the specific sample under study (column 1), the full sample of college educated women, as well as the full sample of women.

Table A5: Differences in average characteristics between mothers and childless women, among women who experienced a miscarriage

	(1) Not childless	(2) Childless	[(1)-(2)] Difference
Age in 2007	50.60	52.08	-1.478***
	(4.478)	(4.721)	(0.117)
Family income in 2007, 1000s SEK	5562.9	4864.6	698.3***
	(4189.9)	(3302.9)	(179.9)
Labor income in 2007, 1000s SEK	2143.2	2108.8	34.45
	(1673.9)	(1779.2)	(43.60)
Compulsory schooling	0.127	0.125	0.00161
	(0.333)	(0.331)	(0.00864)
High school	0.450	0.418	0.0323*
	(0.498)	(0.493)	(0.0129)
College	0.421	0.453	-0.0326*
	(0.494)	(0.498)	(0.0128)
Married	0.590	0.377	0.214***
	(0.492)	(0.485)	(0.0128)
Divorced	0.223	0.182	0.0409***
	(0.416)	(0.386)	(0.0108)
Never married	0.170	0.428	-0.258***
	(0.375)	(0.495)	(0.00990)
Widowed	0.0168	0.0123	0.00455
	(0.129)	(0.110)	(0.00332)
Same-sex partnership	0.0000860	0.00129	-0.00121***
	(0.00927)	(0.0359)	(0.000304)
Observations	34893	1548	

NOTES.— Means, standard deviations and differences in mean characteristics. The sample consists of mothers aged 45 or older in 2007 and who experienced a miscarriage sometime between 1987 and 2005. *p<0.1, **p<0.05 ***p<0.01.

Table A6: Summary statistics for women included in the analysis of the effect of birth spacing on earnings

	Mean
Miscarriage	0.0242
	(0.154)
Age at 1st birth	29.32
	(3.627)
Years between child 1 and child 2 (birth spacing)	2.840
	(1.525)
Non-Nordic background	0.0946
	(0.293)
Age at labor market entry	25.61
	(3.215)
Live in a large city (pre 1st birth)	0.278
	(0.448)
Pre-birth labor income (SEK)	212635.7
	(111590.8)
Pre-birth monthly wage (SEK)	20428.1
	(6151.5)
Observations	208800

NOTES.— The table reports means and standard deviations in parentheses.

Table A7: Falsification test: Reduced form estimates of the effect of miscarriage between the first and second live birth on pre (first) birth income

	Income, SEK OLS	Observations
Outcome measured at		
Year of first birth	0.418	200051
	(1.222)	
Year of first birth -1	1.874	193451
	(1.563)	
Year of first birth -2	0.646	183931
	(1.400)	
Year of first birth -3	-0.831	174922
	(1.506)	
Year of first birth -4	-0.916	166118
	(1.467)	
Year of first birth -5	-2.057	156836
	(1.446)	

NOTES.— Included covariates are the number of pre-natal hospitalizations, a dummy for non-nordic background, dummies for high school and college, a full set of dummies for age at first birth and a full set of dummies for cohort. *p<0.1, **p<0.05 ***p<0.01.

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