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THE EFFECT OF CHILDREN ON EARNINGS USING EXOGENOUS VARIATION IN FAMILY SIZE: SWEDISH EVIDENCE

by

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The Effect of Children on Earnings Using Exogenous Variation in Family Size: Swedish Evidence*

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Abstract

This paper takes advantage of an exogeneous variation in the sex composition of previous children, to study the effect of an additional child on women's earnings. I use OLS and IV as well as quantile regression to analyze the impact of an increase in family size on labour force participation and level of earnings from 1980-2005 Swedish register data. The IV technique produces estimates that are not systematically different from those from OLS, at the expense of a low precision. Including men in the analysis shows that fathers' labour force outcomes are less likely to be affected by an increase in family size compared to mothers. My findings indicate that having an additional child has a stronger negative impact on earnings than on labour force participation. However, there is evidence of catching-up effect over time, as women tend to recover gradually from the negative earnings effect. Using different time perspective, the results remain stable with respect to the rapid expansion of the Swedish family policies. The quantile regression approach suggests that other mechanisms than childbearing lie behind the large wage gap at the top of the wage distribution, often referred to, in Sweden, as the glass ceiling pattern.

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1 Introduction

The gender wage gap has been persistent and substantial over the past decades in most countries. In Sweden, the official statistics indicate that, on average, women's wages are about 18 percent lower than men's across the whole labour market. Taking into account different factors such as occupation, age, working time, education and sector, puts the overall wage differential at 8 percent. (SCB, 2008) This may seem small in comparison to the pay differential between women and men in other countries, but this figure is conspicuously important, considering the small wage dispersion in Sweden. Likewise, the gender gap is particularly large in the top of the Swedish income distribution. Albrecht, Björklund and Vroman (2003) use the term "glass ceiling" effect to describe how the gender wage gap is predominant in the upper tail of the wage distribution, a reflection of the fact that women do well in the labour market up to a point after which there is an effective limit on their prospects.

The wage differential between women and men is often ascribed to the simple fact that motherhood tend to have a negative impact on employed women's wages. Prior studies provide evidence of a motherhood penalty. Among the findings from U.S. data is a wage penalty of 6 percent for mothers with one child and 13 percent if mothers have two or more children (Waldfogel, 1997). Also Budig and England (2001) report a motherhood penalty of 7 percent per child, which narrows down to 5 percent after controlling for experience.

However, the studies based on Swedish data find little effect when looking at the impact of children on women's labour market outcomes. Albrecht, Sundström and Vroman (1999) show that a year of formal parental leave has a small, but statistically negative effect on Swedish women's wages. Also, Harkness and Waldfogel (1999) find 6 percent earnings penalty for two children, and 10 percent penalty for two or more, when using data for Swedish women who worked full-time, in a cross-country

comparison.

Studies of the effect of children on women's wages have been fewer in the Nordic countries. Estimating the causal effect of fertility on women's wages is complicated due to the endogeneity problem that arises, when simply using number of children as explanatory variable in an OLS equation. Obviously, it would cost more to have an additional child if the individual has good career prospects and a possibility of advancement in her career. In this sense, women with kids are not likely to have as good job opportunities as women who do not have children. A source of exogenous variation is therefore necessary in order to identify a causal relationship between fertility and wages. The present paper begins to fill this paucity in the literature by estimating the causal effect of having one more child on Swedish women's earnings. Moreover, this study takes the analysis a step further by using the sex-mix of previous offspring as an exogenous variation in family size, following the identification strategy in Angrist and Evans (1998).

It has long been recognized that gender preferences have significant implication on the number of children born. In Sweden, Andersson, Hank and Vikat (2007) show that, two-child families with unbalanced sex composition of children have a distinct preference for a third child of the missing sex. This parental preference for mixed sibling-sex composition, which tends to influence them to have a third child, offers a potential source of exogenous variation in the estimation of the effect of children on women's labour supply and wages. Angrist and Evans (1998) take advantage of the sex-mix of previous children and twin births in U.S. data to estimate the effect of fertility on parents' labour supply. Their study confirms that children tend to reduce the amount of labour supplied to the market by their mothers. Using sex-mix instrument, they find no labour supply effects among college-educated women and women of high earnings husbands. Their results also show very little response to

changes in family size in the labour market behavior of married men.

Among the studies adopting the same approach in applying an exogenous variation of family size on female labour supply outcomes, as in Angrist and Evans (1998), include Iacovou (2001) for the U.K., Chun and Oh (2002) for Korea, Maurin and Moschion (2006) for France, Ebenstein (2007) for Taiwan, and a recent papper by Daouli, Demoussis and Giannakopoulos (2008) for Greece. Also, Michaud and Tatsiramos (2008) investigate the direct and long-run effects of fertility on female employment in seven European countries, excluding Sweden.

The objective of this paper is, primarily, to estimate the causal effect of having one more child on Swedish women's earnings. Following the identification strategy in Angrist and Evans (1998), I use the sex-mix of previous offspring as an exogenous variation in family size. Studying the sex-mix of two previous children implies that I need to observe individuals with two or more children, that is, those who, at least, had two career interruptions in different periods following childbearing. Meanwhile, using twin birth as an instrument for the number of children like Angrist and Evans (1998) have done, can be argued to be less representative since it only involves one disruption in labour market activity. For my purposes, working with the sex-mix of previous children is preferred over the use of twin birth for the reason that multiple career breaks are likely to have more impact on labour supply and earnings.

This study makes several contributions to the literature. First, I compare the results of the impact of one more child on women's earnings from the 2SLS to those from the OLS. Second, since still few studies consider the effect of children on men's labour force outcomes, I examine how the mentioned effects differ across gender by including men in the analysis.

Third, unlike the above studies, I extend the analysis by separating the effect of having one more child on parental earnings from that on labour force participation,

conditional on having a job. This practice helps capture the total effect on earnings, at the same time as it includes non-participants in the analysis.

Fourth, I consider the contrast between the short-run and long-run effect of an increase in family size. Considering different time perspective is relevant from an economic point of view since long time-out of employment after motherhood may have consequences for long-term career outcomes and the sustainability of the pension systems. I use high-quality data set, which has advantages over the data used by Angrist and Evans (1998). The main advantage is that I am able to follow up the development of individuals' earnings during a 25 years period beginning from 1980 until 2005. This study is, to my knowledge, the first in Swedish data, which covers such a lengthy period of time and gives a preview of the long- run consequence of childbearing on women's career and their lifetime earnings.

Fifth, I use a quantile regression technique to explore which part of the earnings distribution is most affected by an increase in family size. This is expected to give an insight into whether having an additional child can explain why women seem to be impeded to advance on the career path to the top of the occupational hierarchy, known as the glass ceiling pattern in Sweden.

And sixth, taking advantage of the long-run aspect of my data set, I inspect whether the effect is stable over time with respect to the rapid expansion of the family policies in Sweden. This part of the analysis is relevant since the Swedish experience is an especially interesting case from the perspective of the target of unconditional gender equality of opportunities, the generous family-friendly policies associated with the historical development of the Swedish welfare state.

The remainder of this study is outlined as follows. Section 2 presents the theory and econometric framework. I describe the data in section 3. My results are reported in section 4: first the comparison between the OLS and 2SLS, then the results from

the quantile regressions. This is followed by an evaluation of the results with respect to changes in the family policies, and the sensitivity analyses. Section 5 concludes.

2 Theory and econometric framework

2.1 The effect of children on women's wages

There are many reasons why children might affect women's labour market outcomes. Children may affect women's wages directly by lowering a woman's effort on the job. Theory predicts that mothers work less to the extent that their wages are lower, their cost of working outside the home is higher, due to childcare costs, and the value of their time in home production is higher. The fall in mothers' labour supply can be attributed to the increased value of women's home time after having a child (Becker, 1985) and the decline in wage rates to a fall in productivity due to reduced time and effort on the job.

Having children might also influence women's wages indirectly by lowering the amount of work experience and tenure accumulated over time (Mincer and Polacheck, 1974). The human capital theory can explain why women with children earn less than their counterparts.[1] Becker (1985) points out that women overall generally have lower wages than men because they have lower levels of wage-enhancing human capital such as education, training, work experience and job tenure. Mothers tend to earn lower wages because they are more likely to work part-time and for a new employer, which means that they have less work experience and tenure in the long-run. Put differently, women with children are paid less as a result of their choices for more flexible work arrangements.

[1] Alternative explanations include differences of occupations of women with children and discrimination against mothers.

2.2 Parental gender preferences and fertility

The sex of previous children has long been known to have an effect on parity progression. One way of analyzing gender preferences is to observe the increase in family size. If parents prefer children of mixed sex, then families endowed with two boys or two girls are more likely to progress to higher parity than those having a boy and a girl. As stated by Ben-Porath and Welch (1976), the fact that parents care about the sex of their children is established by showing dependence of the tendency to have more children and the sex composition of earlier children.

The sex of women's previous children has an effect on their subsequent childbearing, and the persistence of that effect among women with two children motivate the use of sex of previous children as a predicting variable to women's fertility (Sloane and Lee 1983). A more recent study by Andersson, Hank and Vikat (2007) finds distinct preferences for at least one child of each sex among parents of two children in Sweden, but no effect of the sex of the firstborn child on second-birth risks.[2]

2.3 Econometric framework

Empirically, estimating the effect of having an additional child on women's wages is not straightforward. Including the number of children as an explanatory variable in an OLS equation would produce biased and inconsistent estimates. An instrumental-variable technique addresses this endogeneity, and offers a way to achieve consistent estimates of the effect of fertility on women's earnings. Like Angrist and Evans (1998), I exploit the exogeneous variation from the sex-mix of previous children, to estimate the effect of having one more child on women's earnings.[3] Parental

[2] Also, according to Bulatao (1981), specific sex preferences are found to be most prominent at the third birth.

[3] This section follows closely Angrist and Evans (1998).

preference for mixed sibling-sex composition, which tends to influence them to have a third child, has an impact on the number of children, but is supposed to be unrelated to any unobserved factors that might affect women's labour supply and earnings. In other words, the sex-mix of previous children, which is, in essence random, is highly correlated with fertility while it is not associated with the error term in the regression equation. A dummy variable *Same sex* indicating whether the sex of the first child is the same as that of the second child is used to predict further childbearing among women who have at least two children.

This regression equation links the endogeneous fertility measure x_i with an indicator of labour-force career y_i:

$$y_i = \alpha + \beta x_i + \varepsilon_i \tag{1}$$

While the IV estimation from (1) only includes one explanatory, the next step is to add a list of exogeneous variables in the equation with the aim of achieving more precise estimates. 2SLS also offers an opportunity to control for possible additive effects of the child sex on mothers' labour force attachment. Since the variable *Same sex* is an interaction term of the sex of the first two children, it is clearly correlated with the sex of the first-born as well as that of the second child. This correlation is likely to bias the results in case the sex of either child influences parents' labour market behavior for other reason than family size. This could be the case if parents would treat their child differently depending on the child's sex, or whether the father's committment to the family is contingent on the sex of the child as in certain Asian cultures where there is a strong preference for sons. Including the sex of the first-born and the second child, *Boy 1st* and *Boy 2nd,* in the equation helps minimize any omitted-variable bias.

Because *Same sex* can be decomposed into two instruments: *Two boys* and *Two girls*, adopting 2SLS method allows for an overidentified model with one endogeneous variable but two instruments. It is then possible to apply an instrument-error overidentification test for 2SLS with both instruments. This would give an indication for whether the *Two boys* and *Two girl*s instruments give the same estimate when used separately.

The regression equation which links fertility to the indicator of labour market career becomes

$$y_i = \alpha_0^{'}\mathbf{w}_i + \alpha_1 s_{1i} + \alpha_2 s_{2i} + \beta x_i + \varepsilon_i \tag{2}$$

where $\mathbf{w}_i$ is a vector of demographic variables such as mother's age, age at first birth and education; s_{1i} and s_{2i} are indicators for the sex of the first-born and the second child of mother i, represented by *Boy 1st* and *Boy 2nd*.

The first-stage equation for the just-identified model which includes only one instrument, *Same sex* is

$$x_i = \pi_0^{'}\mathbf{w}_i + \pi_1 s_{1i} + \pi_2 s_{2i} + \gamma(Same\ sex) + \eta_i \tag{3}$$

The overidentified specification which includes the two components of *Same sex*, that is, *Two boys* and *Two girls* can be written:

$$y_i = \alpha_0^{'}\mathbf{w}_i + \alpha_1 s_{1i} + \beta x_i + \varepsilon_i \tag{4}$$

One of s_{1i} and s_{2i} is dropped from equation (5) because s_{1i}, s_{2i} and *Two boys* (s_{1i} s_{2i}) as well as *Two girls*$[(1-s_1)(1-s_2)]$ are linearly dependent. The alternative first-stage equation is

$$x_i = \pi_0^{'}\mathbf{w}_i + \pi_1 s_{1i} + \gamma_0(Two\ boys) + \gamma_1(Two\ girls) + \eta_i \tag{5}$$

2.3.1 Total effects on earnings

There is a concern about the selection bias that arises invariably when using earnings as a measure of labour force attachment in the estimation. It is no exception here since there are individuals who participate in the labour market and those who do not. A way to also include non-participants in the analysis, is to estimate two different models. First, the probability of participating in the labour market $P(Y > 0)$ is estimated by a linear probability model (LPM) with a dummy variable equal to one if the individual has earnings greater than zero. The second component $E(Y|Y > 0)$ represents the expected value of earnings given that individual has a positive earnings, and is estimated by OLS with the log of earnings as dependent variable. Putting together these two components captures the total effect on earnings:

$$E(Y) = P(Y > 0)E(Y|Y > 0) \tag{6}$$

All in all, three different specifications are estimated for each model. The first regression is a simple OLS equation (3) relating the number of children to labour earnings. The second is a 2SLS applied to the same equation using *Same sex* as an instrument for the endogenous fertility variable. As a sensitivity analysis, a third specification where the instrument *Same sex* is replaced by *Two boys* and *Two girls* resumed in equation (5) is also estimated.

3 Data

I use a high-quality Swedish data set, which stems from population registers at Statistics Sweden (SCB). Initially, the data set consists of a 35 percent random sample of Sweden-born individuals. The sample is limited to women who were 23-35 years old in 1980, and men aged 23-40 in 1980.[4] This age restriction ensures, among other things, that women included in the sample are of childbearing age. Since the instrument variable is based on a comparison of the sex of parents' previous children, it is necessary to include individuals with two or more children to be included in the sample. The individuals in the sample are given time to complete the transition to third birth within a few years before or at the same year I observe their earnings. Another criterion implemented, therefore, is that their second child is born between 1965 and 1980.

Descriptive statistics for the sample of women and men are shown in Table 1. The final sample is made up of 103,966 women and 119,976 men who have two or more children at the end of 1980, whose first child is less than 18 years old and second child was born during the period of 1965-1980.[5] The mean age in 1980 is 30.8 for women, and 34.1 for men. On average, women become mothers at an early age of 22.2 compared to 24.8 for fathers. The fertility and demographic variables by *Same sex* show that women with previous offspring of the same sex tend to have more children and are more likely to have two boys than two girls. The corresponding sample statistics for men tell almost the same story.

Since this paper strives to establish a meaningful causal relationship between fertility and labour earnings variable, it is of primary importance that the instrument variable *Same sex*, as much as possible, closely resembles a random assignment. Al-

[4] Men and women included in the separate samples are not necessarily married or living together.

[5] The later restriction is to insure that the first child has not moved from home yet at the time of the study.

though the sex of a child is obviously random, a basic attempt to avoid any concerns is to examine the difference in mean for a few demographic variables of the women with previous children of the same sex and those who have children of mixed sex in the sample. Examples of such variables are *Age in 1980*, *Age at first birth* and *Years of education*. For this purpose, Table 1 indicates that mothers (fathers) having children of same sex and mothers (fathers) with children of mixed sex, have similar characteristics in demographic variables before the treatment, that is, the arrival of a third child. Both groups have the same age in 1980, same age at first birth and number of years of schooling.

How much does the sex composition of previous offspring influence progression to higher parity? Table 2 presents the sample characteristics of women (men) with one boy and one girl, those for two girls are in the second row followed by those for two boys in the third row. The following rows display corresponding figures for women with two children of same sex respective mixed sex. The figures show the fact that women and men with two children of the same sex are more likely to have a third child than those with one boy and one girl. Half of the women and men in the sample have two first children of the same sex and about 48.1 percent of women versus 47.9 percent of men proceed on having a third child. The corresponding fraction for those with one boy and one girl is slightly lower, 42.4 percent for women and 41.9 percent for men. The difference between the same-sex and mixed-sex group average is 0.057 versus 0.060, hence, at most 5.7 (6) percent of Swedish women (men) have an additional child as a result of preference for children of mixed sex.

The figures in the third and fourth columns, third row, allude to the fact that mothers of two boys are more likely to have a third child than mothers of two girls. This feature of the data is consistent with what Andersson, Hank and Vikat (2007) have observed, mainly that Swedish parents seem to develop a preference of having

a girl for third births.

The data set contains information about the income of all individuals in the sample. I use annual labour earnings to represent women's and men's wages. Earnings include income from work, wages and salaries; besides, these cover self-employment, sickness benefits and parents' allowance. Using this particular measure of earnings is suitable for the purpose of this study, since labour earnings reflect both hourly wages and labour supply. Their labour earnings is measured over 25 years, beginning from 1980 until 2005, a period that is long enough to enable a look at the long-term impact of an increase in family size on women's and men's career. Likewise, it allows a comparison of lifetime earnings of women and men who have two or more children to those who did not have a third child.

There are two main outcome variables in the analysis: the first one is a dummy variable indicating labour force participation, which equals one if individual has positive earnings, and the second one is yearly log labour earnings. Table 3 reports the fraction of the individuals in the sample who have positive earnings during the period of 1980-2005.[6] In 1980, 89 percent of the women in the sample have earnings greater than zero. Hereafter, I define labour force participation as the fact that an individual has earnings greater than zero. This proportion increases to 93 percent in 1990, and further to 97 percent in 1995. The fraction of women who are active in the labour market falls slightly to 94 percent in 1995 and in the later years, this fraction goes down even more in 2000 and 2005 as a share of the women in the sample are going towards retirement. There is not much variation when it comes to the proportion of men who are employed and have positive earnings. The fraction of the sample, shown in the lower part of Table 3, is quite stable over the years and lies around 99 percent for 1980-1990 to about 90 percent in the later years, when presumably some

[6]Individuals with missing earnings values are dropped to exclude from the sample those who have died or moved abroad.

have retired.

The independent variable of interest, which represents the number of children, is *Third Child*, a dummy variable which takes the value of one if individual has a third child before or during the year her labour earnings are observed. For instance, the variable *Third Child* 1990 is equal to one if the individual has a third child born in 1990 or before. Otherwise *Third Child* 1990 equals zero if individual has only two children or her third child is born after 1990. Note that year 1980 is comparable to a reference year as the analysis starts then and *Third Child* 1980 is equal to one if the individual has a third child born in 1980. Table 4 presents the fraction of women and men in the sample, who have a third child before or at the time their earnings are observed. The share of women in the sample whose third child was born during 1980 is 25 percent, the corresponding figure is 26 percent for men. This share almost doubles in 1985 with 40 percent for both women and men, and stays stable in the later years, around 45 percent for women while roughly 50 percent for men.

4 Results

I start by reporting the results from a linear probability model of labour force participation, estimated by OLS. The estimates for women presented in the upper part of Table 5 show that having one more child has a negative effect on labour force participation for mothers, the effect being highly significant. Looking at the first row, a third birth in 1980 tends to reduce the probability for a mother to participate in the labour market by 5 percentage points. This negative effect goes down to roughly 2 percentage points from 1990 to 2000, to further decrease to 1.5 percentage points in 2005. Accordingly, the negative labour supply effect persists but becomes smaller when the child grows up. Harkness and Waldfogel (1999) also found that in Sweden, women with infants and pre-school age children are significantly less likely to be

employed, compared to women with school age children.

When applying an instrumental variable technique and using *Same sex* as an instrument for the variable *Third Child*, the women's results reveal a different story. At first sight, the 2SLS results seem to indicate that having one more child does not have a negative impact on mothers' labour force supply. However, at a closer look, most of the estimates are positive and not statistically significant, except for 1981 to 1985. The presence of a third child lowers the probability for a mother to participate in the labour market by 9 percentage points in 1981 and 1982. The decrease in probability of being active in the labour market is 11 percentage points in 1983 and 7 percentage points in 1984 and 1985, when using the sex of the mother's two previous children as an instrument for a third birth. This negative impact on labour participation is larger than the corresponding effects from the OLS, which are 6 and 5 percentage points in 1984 and 1985. However, the magnitude of the effect from the IV estimation is about the same as that from OLS for the years 1981 to 1983.

Turning to the results for men in the lower part of Table 5, the OLS estimates display, essentially, a significant negative effect of an increase of family size on men's labour supply. The size of the OLS estimates for men is much smaller than for women. Having a third child reduces the probability for fathers to be out in the labour market by 0.1 percentage points in 1980. The negative effect on fathers' labour supply is 0.5 percentage points in 1995 and 2000. While the effect of a third birth on mothers' labour force participation decreases gradually in the long-run, the effect on fathers's supply seems quite stable over the years. This, it seems, contradicts past results which suggest that young children are associated with longer work hours for men (Pencavel, 1986).

The 2SLS estimates for men turn positive for most of the years when using *Same*

sex as an instrument for a third birth. However, these effects fail to be statistically significant. This suggests that fathers' labour participation is far less likely to be affected by an increase in family size compared to mothers. Also, Angrist & Evans (1998) have found very little response to changes in family size in the labour market behaviour of husbands in their sample when using the same estimation technique and instrument. This leads to believe that even when more children usually require more time spent by fathers caring for them, this is done at the expense of his leisure time rather than his working hours.

The second part of the empirical analysis deals with the expected value of earnings given that the individual has positive earnings. The outcome variable here is the log labour earnings and the main explanatory variable is *Third Child* as above.

Starting from the results for women in the first rows of Table 6, the OLS estimates display the same negative effect as before, only now, the coefficients tend to be much larger than those from Table 5. Results from the OLS imply that a third birth has, overall, a significant negative effect on women's earnings. The negative impact of having a third child on earnings during 1980 is about 40 percent. This figure may seem quite large and one has to be careful in interpreting it, keeping in mind that year 1980 is similar to a reference year for the individual's earnings. While the negative earnings effect is 35 percent in 1985, it decreases to 18 percent in 1990, and continues to diminish gradually in magnitude following the years. In 2005, the presence of a third child reduces women's earnings by only 4 percent, suggesting a possible catch-up effect.

Looking at the 2SLS results of the effect of third birth on women's earnings in the second row of Table 6, using *Same sex* as an instrument for having one more child produces coefficients estimates that are negative and smaller in magnitude. The presence of a third child during the year when mothers' earnings are observed, in

1985, implies a negative effect of 32 percent on earnings. As with the OLS results, the magnitude of the 2SLS coefficients also decreases gradually following the years, although the 2SLS estimates from 1990 to 2005 are not statistically significant. The 2SLS results give some evidence of a rebound effect on earnings after childbearing when using the sex of the mother's two previous child as an instrument for a third birth.

When it comes to the pay effects of one more child on fathers' earnings in the lower part of Table 6, the OLS results depict significantly negative estimates. Again, the coefficients estimates for men are much smaller than those for women. The presence of a third child leads to an earnings reduction of around 4.5 to 5 percent from 1985 to the year 2000. As for the results when applying an IV technique, the coefficients estimates for men are smaller, mostly negative, and of poor precision. This suggests that men's earnings are less sensitive to an increase in family size compared to those of mothers. In fact, a previous study evaluating the magnitude of family gap across countries, reports that men with two or more children tend to earn more (Waldfogel, 1998).

Now, some additional comments about the results. For lack of space, only results from five years intervals are presented in the tables, however, Figure 1 and 2 provide an intuition for the yearly fluctuations of the estimates results. The figures depict more dramatic variation in the effect of a third birth on women's labour outcomes. Also, they show that fathers' labour market career remains fairly unaffected by an increase in family size compared to mothers, however, there seems to be a trend of convergence towards more constant effect in the later years.

The F-statistics from the first-stage estimation reported underneath the point estimates of the results in Table 5 and 6, are considerably large and above 200. Their magnitude is well beyond the rule-of-thumb of 10 (Staiger and Stock, 1997). Also

Stock and Yogo (2005) suggest a critical value of 16.38 for the first-stage F-statistic for a single endogenous regressor and one instrumental variable. Therefore, the sizes of the first-stage F-statistics are large enough to elude any concerns about the predictive power of the instrument.

Overall, the 2SLS estimates are not systematically different from its OLS counterparts. A Hausman test (Hausman, 1978), which evaluates the difference between the estimates from two different methods reveals, not surprisingly, that the 2SLS and the OLS estimates are not statistically different from each other. In fact, the p-value of the Hausman test for each and every year, is around one, implying that the null hypothesis that coefficients from 2SLS and OLS are the same, cannot be rejected.

Moreover, as can be seen throughout the analysis, the standard errors of the 2SLS estimates are much larger than those of the OLS, and the precision is consequently lower. Although the low precision may seem to be a disappointing lack of effect in my study, given that the main purpose is to use the exogenous variation from the sex-mix of the two previous children to identify a causal effect, it is usually the price to pay when using the IV technique. Nonetheless, the patterns emerging from the results provide important insights. For instance, the results depict a clear difference depending on the time perspective. Considering the contrast between the short and the long-run earnings impact of a third birth, the short-term child penalty is more pronounced than the long-term effects.

Likewise, looking at the long-term effects of third birth establishes the presence of a catching up effects in women's earnings, which is another interesting feature of my results. This is even the case when looking at the impact of having a third child on earnings using the IV technique. Also, previous research has found a rebound effect, when considering the effects of work career interruptions, which for women are primarily due to childbearing. Mincer and Ofek (1982) and Corcoran, Duncan

and Ponza (1983) discern that real wages at reentry are, on the average, lower than at the point of the labour market withdrawal, but a return to work is followed by a relatively rapid growth in wages. Mincer and Ofek (1982) attribute this wage growth upon return to the labour market, to a restoration of human capital associated with accumulation of job tenure.

4.1 Quantile regression results

In this section, I explore further the effect of an additional child on parents' wages by using a quantile regression estimation. Quantile regression is an adequate tool when measuring whether the effect of having an additional child differs across the whole earnings distribution of the parents, and yet account for typical control variables such as age at first birth and cohort dummies.[7]It is commonly used when there is a need to know the effect that covariates have in certain aspects of the response distribution. Moreover, this part of the analysis can also give an insight into whether having an additional child could explain the "glass ceiling" pattern in Sweden.

I use the same sample as before, the results of the quantile regression estimation are in Table 7. The immediate emerging pattern to discern, is the overall large size of the estimates at the bottom of the log earnings distribution, for both women and men. The coefficient estimates appear to decrease the higher the quantiles are. In the upper part of Table 7, the quantile regression estimates for women display particularly large negative effect of children on mothers' earnings, at the bottom percentiles compared to the top of the log earnings distribution. The lower in the earnings category a mother belongs, the more affected her earnings are of a family increase. This can be due to either a wage penalty of having one more child, or an indication of a negative labour

[7]Note that a quantile regression technique, which takes into account the IV approach as in Abadie, Angrist and Imbens (2002) is feasible here. However, this aspect is not considered in the analysis given that the IV estimates are quite close to those from OLS.

supply effect, from a decrease in hours worked following childbearing. The negative effect on earnings is notably more pronounced for the bottom quantile of year 1980 and 1985. The difference between the coefficients at the top and the lower end of mothers'log earnings distribution is especially large for the reference year 1980. For all women in the sample, having a third child means reduced earnings in all categories of the log earnings distribution.

Again, there is evidence of earnings rebound as the size of the coefficients gets less negative at the lower percentiles of year 2000 and 2005. In essence, the effect of having one more child on mothers' log earnings decreases successively following the quantiles, and the length of the time after the third birth.

Now to the results for men in the lower part of Table 7. As is the case for women, the quantile estimates also are large at the 0.05 quantile and decreases towards the 0.90. However, the difference in magnitude is not as dramatic as in the women's results. The largest effect at the bottom of the log distribution is for year 1985, 1990 and 1995. Nonetheless, the negative impact of having a third child on fathers' earnings turns positive for fathers at the top of the log earnings distribution. The coefficients estimates are of a positive sign for the 0.90 and 0.95 quantiles in 2000 and 2005. This reverse sign of the effect implies that fathers at the top of the earnings distribution tend to be better off by an increase in family size, than those who lie at the bottom.

Overall, the median estimate seems lower compared to those from the OLS. The quantile regression delivers, however, precise and highly significant estimates.

Albrecht, Björklund and Vroman (2003) use the term "glass ceiling" effect to describe how the gender wage gap is predominant in the upper tail, a reflection of the fact that women do well in the labour market up to a point after which there is an effective limit on their prospects. Using quantile regression technique on 1993 Swedish

data, they find that the gap was increasing across the wages distribution, and was particularly large from the 85th percentile to the very top of the wage distribution.

Here, the quantile regression results show that parents who lie at the bottom of the earnings distribution are more penalized by the presence of a third child. In other words, the negative impact of having one more child is larger at the lower quantiles of the distribution. One way to explain my results is the fact that an important share of parents with younger children, more often mothers, are working part-time. Actually, about 40 percent of mothers work part-time in Sweden. Accordingly, my study gives no evidence to suggest that having an additional child can explain the large wage gap at the top of the wage distribution. Although I cannot rule it out as a possible underlying factor of the glass ceiling pattern in Sweden, other mechanisms than an increase in family size are at work when it comes to understanding the determinant of the large gender gap at the top of the earnings distribution.

4.2 Has the effect of having one more child changed over the years?

It is well known that Sweden is among the Nordic welfare states, where accomodating family policies, and labour market measures facilitate a combination of motherhood and work life for women. A recent survey paper by Datta Gupta, Smith and Verner (2008), evaluates the impact of Nordic countries's family-friendly policies on employment, wages and children. They uncover another side of the flexible social system and labour market measures, mainly that family-friendly scheme can have reverse effects on women's wages and career. Women not only suffer a loss of important skills when being away from labour market activity for a longer period of time, but also are at disadvantage in terms of the gender equality status.

Another prospect worth developing is an analysis of how different family policies,

at a different point in time, could explain the change in the results in the long-run. Women's situation has conspicuously changed since 1980 and there has been an increase in female labour force participation and full-time job. It is reasonable to argue that different reform such as child care provision, extended parental leave and increased family allowances are likely to affect the impact of an increase in family size on labour outcomes. In this section, my purpose is to inspect to what extent the various family and labour market policies influence women's decision to participate in the labour market and to have children. For more detailed survey of the development of family policy in Sweden, see Ferrarini (2006).

To see whether the effect found in my results is stable over time with respect to rapid expansion of the family policies in Sweden, I select a new sample of women who are 23-35 years old in 1995, and men aged 23-40 in 1995. I use exactly the same sample restrictions as above, only now I start out with individuals whose second child was born between 1980 and 1995, instead of 1965 to 1980. Thereafter, I follow up on their labour earnings, during a 10 years period beginning from 1995 until 2005. The results are presented in Table 8 and 9. In general, there is no prominent differences between the results in Table 5 and those displayed in Table 8. For women, the presence of a third child still has a negative effect on labour force participation. There has been some changes in the sign of the coefficient estimates, but these changes are not significant. Comparing the effect of one more child on log earnings in Table 6 to Table 9, the only noticeable difference is that IV estimates for year 2000 is now negative and statistically significant. As for the rest, there is still an indication that in the long-run, mothers' earnings recover successively after childbearing although the estimates from the new samples are smaller.

As for men, there is basically no change when looking at the effect of one more children on fathers' labour supply. Moreover, Table 6 and Table 9 show that the effect

of an increase in family size is about the same magnitude, despite different changes in time. Again, fathers' labour market behaviour stays unresponsive to change in family size.

All in all, the results are quite stable despite changes in family policies in Sweden over the years, and diverse family-friendly labour market measures seem not drive the long-term pattern in the results.

4.3 Sensitivity analysis

A couple of sensitivity analyses are conducted with the aim of testing the robustness and improving the precision of my results. The first one involves the estimation of a third specification where the instrument *Same sex* is replaced by *Two boys* and *Two girls* instruments, as shown in equation (5) in section 2. The main advantage of decomposing *Same sex* into *Two boys* and *Two girls* is the use of an overidentification test to find out whether using only *Two boys* in the equation would ensue a statistically different result compared to using *Two girls* as an instrument. The p-values of this overidentification test, the Sargan statistics, disclose no sign of difference between the use of *Two boys* and *Two girl*. The exception is the Sargan statistics for year 1980, which is significant at 5 percent level for women and for year 1985, which is marginally significant in men's results. The coefficients estimates from the regression with *Two boys* and *Two girls* do not differ much from those from using *Same sex* as the only instrument. Then separating *Same sex* into two instruments enhances neither the magnitude nor the precision of the estimates. Accordingly, it does not matter which instrument is included in the equation. Angrist and Evans (1998) also encounter similar pattern in their analysis when using two instrument variables instead of one.

Second, after studying the sample of individuals with earnings greater than zero,

next process is to increase the earnings limit to 50,000 SEK. Not surprisingly, this improves the precision of the estimates but decreases the magnitude of the effect, especially for the OLS estimates. The estimates for the influence of an increase in family size on labour supply have not improved much, neither in magnitude nor in precision. As for the rest, the same pattern remains when the income limit is raised to 50,000 SEK.

The third part consists of including the number of years of schooling in 1990 in the regression equation. Detailed information about the offspring's educational attainment from the 1990 version of the Swedish education register[8] is converted into years of schooling and into level of education. Starting with primary school, which is obligatory at 6 or 7 years of age, I define seven levels of education with a corresponding total number of years: primary education corresponds to 7 years, lower secondary education to 9 years, short and long upper secondary education amount to 11 and 12 years, short and long university approximate to 14 and 15.5 years and finally PhD studies amount to 19 years. The estimates of the effect of one more child on labour force participation do not change noticeably much, for both women and men. This is the case for both estimation methods. As for the impact of the presence of a third child on the log earnings, accounting for education, measured as *Years of schooling in 1990*, increased the magnitude of the estimates slightly for both women and men. However, the precision remains the same when using OLS and 2SLS. The same general pattern of results remains present when replacing years of schooling by a set of dummies, which indicate the level of education. Though this is a further step towards a more detailed definition of the education variable, it does not alter the results reported above.

[8]Ideally, information about individual's education from 1980 would be more suitable in this analysis. However, the prior version of education register is not good enough in terms of quality and precision.

5 Conclusions

The objective of this paper is, primarily, to estimate the causal effect of having one more child on Swedish women's earnings. Following the identification strategy in Angrist and Evans (1998), I use the sex-mix of previous offspring as an exogenous variation in family size. This study makes several contributions to the literature.

First, a comparison between the OLS and IV shows that the estimates from both estimation approaches are not systematically different, and the Hausman test indicates that they are not significantly different from each other. However, the OLS delivers statistically significant results, while using Same sex as an instrument for the variable Third Child produce estimates with large standard errors and of poor precision. Although this low precision may seem to be a disappointing result in my study, it is usually the price to pay when using the IV technique.

Second, including men in the analysis reveals that fathers' labour market outcomes indicate very little or no response at all to changes in family size. This lack of effect is more notable for labour force participation, which suggests that even when more children usually require more time spent by fathers caring for them, this is done at the expense of his leisure time rather than his working hours.

Third, when separating the effect of having one more child on earnings from that on labour force participation, conditional on having a job, the impact on earnings tends to be more pronounced compared to that on the supply of labour. While the effect of a third birth on mothers' labour force participation decreases gradually over time, the effect on fathers' labour supply seems quite stable over the years. When it comes to the effect of an additional child on parental earnings, men's earnings are less sensitive to an increase in family size than mothers'.

Fourth, I take advantage of data on individuals' earnings during a 25 years period beginning from 1980 until 2005, to compare the short- and long-term effect of having

one more child on parental earnings. The results depict a clear difference in the extent of the effect depending on the time perspective. The short-run impact of a third birth on earnings is more important than the long-term effect. Also, the long-term effect establishes the presence of catching-up effects on women's earnings. This rebound effect, in labour force participation and earnings, implies that women who return to work after having a third child are likely to recover gradually from the negative earnings effect.

Fifth, using quantile regression to explore the part of parental log earnings distribution that is mostly influenced by an increase in family size shows a substantial effect at the bottom percentiles. In essence, the lower in the earnings category a mother belongs, the more affected her earnings are of a family increase. A possible explanation is the important share of mothers, about 40 percent, with younger children, who are working part-time in Sweden. As for fathers, the negative impact of having a third child on earnings turns positive for fathers at the top of the log earnings distribution. Put differently, fathers at the top tend to be better off by an increase in family size than those who lie at the bottom tail of the earnings distribution. Taken together, the quantile regression results do not confirm that having an additional child can explain the glass ceiling pattern in Sweden.

And sixth, focussing on the long-run aspect of my data set, I inspect whether the effect in my results is stable over time with respect to the rapid expansion of the Swedish family policies. Starting up with a new sample and studying their labour market outcomes 10 years later exposes, as before, a negative effect of having a third child on mothers' labour force participation and earnings. Moreover, the catching-up effect still remains. It follows that different family policies, at a different point in time, do not drive the change in my results in the long-run.

All in all, the results are stable to different changes in specification and variables

included in the analysis. My study gives no evidence to suggest that having an additional child can explain the large wage gap at the top of the wage distribution. Although I cannot rule it out as a possible underlying factor of the glass ceiling pattern in Sweden, other mechanisms are at work when it comes to understanding the determinants of the gender wage gap. Further research is needed to uncover those mechanisms.

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Table 1: Descriptive Statistics for Women Aged 23-35 and Men 23-40 in 1980

Variable	Mean and (standard deviation)					
	Same Sex=0		Same Sex=1		All	
	Women	Men	Women	Men	Women	Men
Number of children	2.59	2.62	2.67	2.70	2.63	2.66
	(0.86)	(0.91)	(0.87)	(0.93)	(0.87)	(0.92)
More than 2 children	0.424	0.419	0.481	0.479	0.453	0.449
(=1 if more than 2 children)	(0.494)	(0.493)	(0.499)	(0.499)	(0.498)	(0.497)
Boy 1st	0.499	0.501	0.530	0.529	0.515	0.515
(=1 if first child was a boy)	(0.500)	(0.500)	(0.499)	(0.499)	(0.500)	(0.500)
Boy 2nd	0.500	0.499	0.530	0.529	0.515	0.514
(=1 if second child was a boy)	(0.500)	(0.500)	(0.499)	(0.499)	(0.500)	(0.500)
Two boys	0	0	0.530	0.529	0.266	0.266
(=1 if first 2 children were boys)	0	0	(0.499)	(0.499)	(0.442)	(0.442)
Two girls	0	0	0.470	0.471	0.236	0.237
(=1 if first 2 children were girls)	0	0	(0.499)	(0.499)	(0.425)	(0.425)
Age in 1980	30.9	34.1	30.9	34.1	30.9	34.1
	(3.2)	(3.7)	(3.2)	(3.7)	(3.2)	(3.7)
Age at first birth	22.1	24.8	22.2	24.8	22.2	24.8
(age when first child was born)	(3.2)	(3.3)	(3.2)	(3.3)	(3.2)	(3.3)
Years of schooling in 1990	11.1	11.0	11.1	11.0	11.13	11.0
	(2.4)	(2.9)	(2.4)	(2.9)	(2.4)	(2.9)
Number of observations	51,736	59,531	52,230	60,445	103,966	119,976

Table 2: Fraction of Sample that Had Another Child by Parity and Sex of Children

Sex of first two children	Fraction of sample		Fraction of sample that had another child	
	Women	Men	Women	Men
one boy, one girl	0.498	0.496	0.424	0.419
			(0.002)	(0.002)
two girls	0.236	0.237	0.476	0.475
			(0.003)	(0.003)
two boys	0.266	0.267	0.486	0.482
			(0.003)	(0.003)
(1) one boy, one girl	0.498	0.497	0.424	0.419
			(0.002)	(0.002)
(2) both same sex	0.502	0.504	0.481	0.479
			(0.002)	(0.002)
difference (2) - (1)			0.057	0.060
			(0.003)	(0.003)
Sample size	103,966	119,976	103,966	119,976

The samples include women aged 23-35 and men 23-40 in 1980, with two or more children and whose second child is born 1965-1980. Standard errors are in parantheses

Table 3: Fraction of Sample with Positive Earnings

	1980	1985	1990	1995	2000	2005
	Women 23-35 in 1975					
Proportion	0.891	0.934	0.968	0.941	0.922	0.860
Sample size	92,640	97,125	100,688	97,815	95,859	89,413
Total sample	103,966					
	Men 23-40 in 1975					
Proportion	0.994	0.990	0.986	0.953	0.927	0.833
Sample size	119,236	118,743	118,311	114,324	111,244	100,000
Total sample	119,976					

The samples consist of all men and women with 2 or more children

Table 4: Fraction of Sample with a Third Child

Dummy var =1 if	1980	1985	1990	1995	2000	2005
third child is born	Women 23-35 in 1975					
Proportion	0.249	0.400	0.445	0.452	0.453	0.453
Sample size	25,850	41,582	46,243	46,993	47,073	47,076
Total sample	103,966					
	Men 23-40 in 1975					
Proportion	0.262	0.395	0.435	0.446	0.448	0.449
Sample size	31,471	47,359	52,206	53,481	53.813	53.908
Total sample	119,976					

The reference group consists of those who have only two children and those who have a third child after the year their earnings is observed

Table 5: OLS and 2SLS Estimates of Labour Force Participation

Dummy var =1 if	1980	1985	1990	1995	2000	2005
labour earnings > 0	Women 23-35 in 1980					
(1) OLS	-0.047	-0.049	-0.022	-0.021	-0.016	-0.015
	(0.002)	(0.017)	(0.001)	(0.001)	(0.002)	(0.002)
(2) 2SLS	0.023	-0.070	0.003	0.027	0.003	0.020
	(0.048)	(0.025)	(0.018)	(0.025)	(0.028)	(0.037)
(2)1st-stage F-statistics	252.17	406.96	381.18	373.36	371.08	370.34
Sample size	103,966	103,966	103,966	103,966	103,966	103,966
	Men 23-40 in 1980					
(1) OLS	-0.001	-0.001	-0.0039	-0.005	-0.005	0.004
	(0.0006)	(0.0006)	(0.0007)	(0.001)	(0.001)	(0.002)
(2) 2SLS	0.001	-0.013	0.006	-0.017	0.016	0.008
	(0.001)	(0.010)	(0.011)	(0.020)	(0.025)	(0.034)
(2)1st-stage F-statistics	261.09	460.92	476.30	456.90	456.98	460.11
Sample size	119,976	119,976	119,976	119,976	119,976	119,976

The table reports coefficient estimates of Third child

In (2): the instrument for Third child is Same sex

Covariates: Age at first birth, Cohort dummies 1945-1957 (women); 1940-1957 (men), Boy 1st, Boy 2nd

Table 6: OLS and 2SLS Estimates of Log Labour Earnings

Log labour earnings	1980	1985	1990	1995	2000	2005
	Women 23-35 in 1980					
(1) OLS	-0.415	-0.327	-0.180	-0.133	-0.072	-0.038
	(0.009)	(0.006)	(0.004)	(0.005)	(0.005)	(0.006)
(2) 2SLS	-0.309	-0.322	-0.092	-0.117	-0.054	-0.079
	(0.186)	(0.098)	(0.072)	(0.083)	(0.083)	(0.099)
(2)1st-stage F-statistics	202.21	364.42	388.74	379.23	359.97	330.02
Sample size	92,640	97,125	100,688	97,815	95,859	89,413
	Men 23-40 in 1980					
(1) OLS	-0.037	-0.047	-0.045	-0.052	-0.037	0.001
	(0.003)	(0.003)	(0.003)	(0.005)	(0.005)	(0.007)
(2) 2SLS	0.013	-0.006	-0.026	0.010	-0.067	-0.089
	(0.074)	(0.058)	(0.056)	(0.081)	(0.085)	(0.107)
(2)1st-stage F-statistics	261.71	450.91	481.16	439.03	437.97	429.66
Sample size	119,236	118,743	118,311	114,324	111,244	100,000

The table reports coefficient estimates of Third child, earnings are in 2005's price level

In (2): the instrument for Third child is Same sex

Covariates: Age at first birth, Cohort dummies 1945-1957 (women); 1940-1957 (men), Boy 1st, Boy 2nd

Table 7: Quantile Regression Estimates of Log Labour Earnings

		Quantile						
Log labour earnings	N	0.05	0.10	0.25	0.50	0.75	0.90	0.95
		Women 23-35 in 1980						
1980	92,640	-0.977	-0.929	-0.723	-0.248	-0.160	-0.097	-0.063
		(0.055)	(0.028)	(0.019)	(0.006)	(0.006)	(0.006)	(0.004)
1985	97,125	-1.090	-0.920	-0.388	-0.190	-0.118	-0.074	-0.056
		(0.041)	(0.022)	(0.009)	(0.004)	(0.003)	(0.004)	(0.004)
1990	100,688	-0.622	-0.386	-0.199	-0.123	-0.081	-0.050	-0.047
		(0.024)	(0.016)	(0.005)	(0.003)	(0.003)	(0.002)	(0.004)
1995	97,815	-0.573	-0.304	-0.128	-0.077	-0.046	-0.041	-0.043
		(0.031)	(0.017)	(0.004)	(0.002)	(0.003)	(0.003)	(0.006)
2000	95,859	-0.309	-0.162	-0.074	-0.042	-0.026	-0.025	-0.024
		(0.025)	(0.015)	(0.005)	(0.002)	(0.003)	(0.003)	(0.007)
2005	89,413	-0.193	-0.091	-0.043	-0.021	-0.012	-0.015	-0.016
		(0.036)	(0.020)	(0.007)	(0.003)	(0.002)	(0.005)	(0.007)
		Men 23-40 in 1980						
1980	119,236	-0.165	-0.141	-0.047	-0.023	-0.005	0.020	0.054
		(0.015)	(0.012)	(0.003)	(0.002)	(0.003)	(0.004)	(0.007)
1985	118,743	-0.263	-0.186	-0.049	-0.023	-0.003	0.024	0.046
		(0.026)	(0.010)	(0.002)	(0.002)	(0.004)	(0.004)	(0.006)
1990	118,311	-0.235	-0.169	-0.045	-0.026	-0.009	0.019	0.039
		(0.013)	(0.100)	(0.002)	(0.002)	(0.003)	(0.006)	(0.008)
1995	114,324	-0.250	-0.202	-0.066	-0.021	-0.003	0.031	0.046
		(0.033)	(0.015)	(0.003)	(0.002)	(0.004)	(0.005)	(0.007)
2000	111,244	-0.207	-0.163	-0.051	-0.017	0.003	0.031	0.059
		(0.022)	(0.016)	(0.004)	(0.002)	(0.003)	(0.005)	(0.006)
2005	100,000	-0.023	-0.037	-0.038	-0.011	0.010	0.046	0.065
		(0.042)	(0.018)	(0.006)	(0.002)	(0.002)	(0.004)	(0.006)

The table reports coefficient estimates of Third child conditional on positive earnings.

Bootstrapped standard errors are in parentheses.

Covariates: Age at first birth, Cohort dummies 1945-1957 (women); 1940-1957 (men), Boy 1st, Boy 2nd

Table 8: OLS and 2SLS Estimates of Labour Force Participation

Dummy var =1 if	1995	2000	2005
labour earnings > 0	Women 23-35 in 1995		
(1) OLS	-0.012	-0.010	-0.012
	(0.002)	(0.002)	(0.002)
(2) 2SLS	-0.012	-0.006	-0.022
	(0.032)	(0.024)	(0.027)
(2)1st-stage F-statistics	208.20	330.79	341.67
Sample size	73,461	73,461	73,461
	Men 23-40 in 1995		
(1) OLS	0.003	0.008	0.001
	(0.001)	(0.001))	(0.021)
(2) 2SLS	-0.026	-0.027	0.007
	(0.021)	(0.017)	(0.021)
(2)1st-stage F-statistics	302.26	414.78	426.68
Sample size	100,136	100,136	100,136

The table reports coefficient estimates of Third child

In (2): the instrument for Third child is Same sex. Covariates: Age at first birth, Cohort dummies 1950-1967 (women); 1955-1967 (men), Boy 1st, Boy 2nd

Table 9: OLS and 2SLS Estimates of Log Labour Earnings

Dummy var =1 if	1995	2000	2005
labour earnings > 0	Women 23-35 in 1995		
(1) OLS	-0.189	-0.199	-0.134
	(0.008)	(0.007)	(0.006)
(2) 2SLS	-0.346	-0.403	-0.113
	(0.150)	(0.103)	(0.090)
(2)1st-stage F-statistics	202.49	322.41	339.93
Sample size	70,415	70,148	68,967
	Men 23-40 in 1995		
(1) OLS	-0.034	-0.022	-0.013
	(0.006)	(0.005)	(0.005)
(2) 2SLS	-0.161	-0.033	-0.042
	(0.112)	(0.076)	(0.073)
(2)1st-stage F-statistics	290.79	404.38	409.35
Sample size	97,452	97,236	95,682

The table reports coefficient estimates of Third child

In (2): the instrument for Third child is Same sex. Covariates: Age at first birth, Cohort dummies 1950-1967 (women); 1955-1967 (men), Boy 1st, Boy 2nd

Figure 1: The Effect of Children on Labour Force Participation

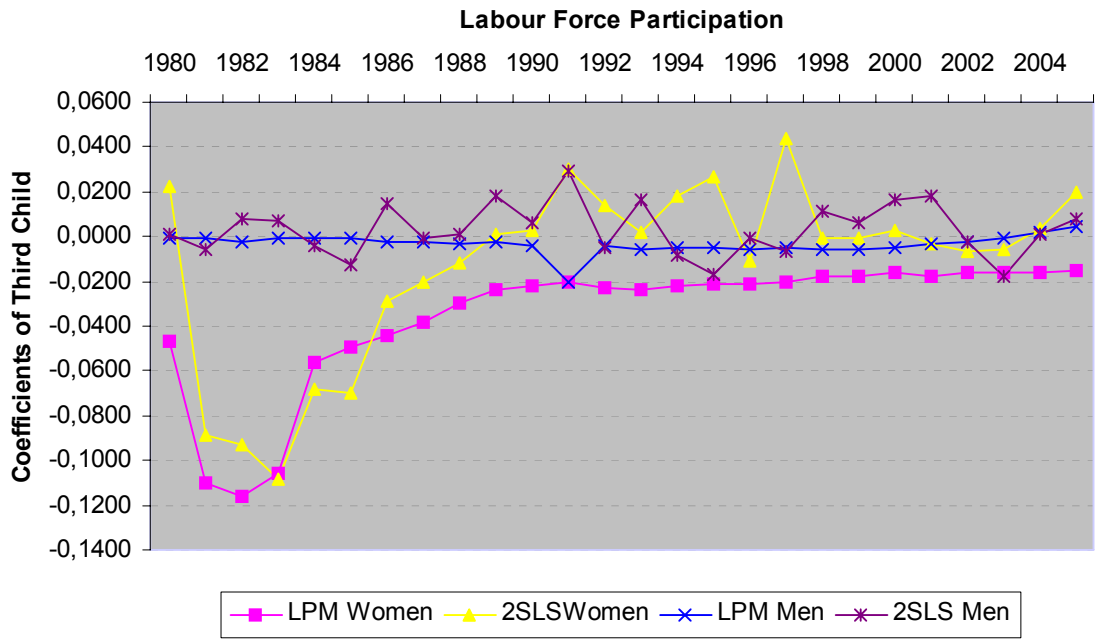
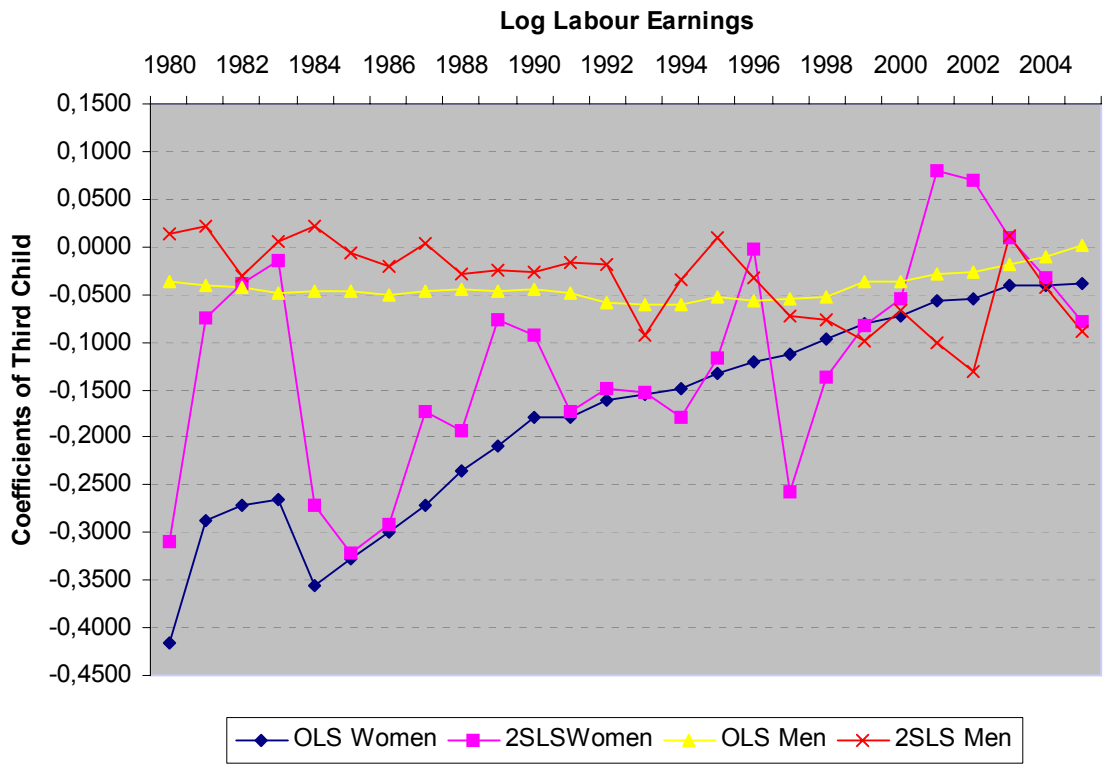


Figure 2: The Effect of Children on Log Earnings



A Appendix

Table A1 : Estimates of Labour Force Participation for Earnings > 50 000 SEK

Dummy var =1 if	1980	1985	1990	1995	2000	2005
labour earnings > 50 000	Women 23-35 in 1980					
(1) OLS	-0.185	-0.146	-0.058	-0.048	-0.029	-0.021
	(0.003)	(0.003)	(0.002)	(0.002)	(0.002)	(0.002)
(2) 2SLS	-0.098	-0.167	-0.016	0.009	-0.018	0.011
	(0.070)	(0.040)	(0.030)	(0.035)	(0.035)	(0.041)
(2)1st-stage F-statistics	252.17	406.93	381.18	373.36	371.08	370.34
Sample size	103,966	103,966	103,966	103,966	103,966	103,966
	Men 23-40 in 1980					
(1) OLS	-0.007	-0.011	-0.010	-0.016	-0.012	0.003
	(0.001)	(0.001)	(0.001)	(0.002)	(0.002)	(0.002)
(2) 2SLS	0.022	-0.003	0.009	-0.012	0.015	0.027
	(0.023)	(0.018)	(0.017)	(0.028)	(0.031)	(0.039)
(2)1st-stage F-statistics	261.09	460.92	476.30	456.90	456.98	460.11
Sample size	119,976	119,976	119,976	119,976	119,976	119,976

The table reports coefficient estimates of Third child

In (2): the instrument for Third child is Same sex

Covariates: Age at first birth, Cohort dummies 1945-1957 (women); 1940-1957 (men), Boy 1st, Boy 2nd

Table A2: Estimates of Log Labour Earnings for Earnings > 50 000 SEK

Log labour earnings	1980	1985	1990	1995	2000	2005
>50 000	Women 23-35 in 1980					
(1) OLS	-0.069	-0.088	-0.095	-0.056	-0.035	-0.015
	(0.004)	(0.003)	(0.002)	(0.002)	(0.003)	(0.003)
(2) 2SLS	0.038	-0.027	-0.055	-0.097	0.025	-0.039
	(0.086)	(0.043)	(0.039)	(0.041)	(0.043)	(0.048)
(2)1st-stage F-statistics	131.58	314.94	374.72	351.68	342.04	331.77
Sample size	72,239	83,479	94,703	91,134	90,541	83,994
	Men 23-40 in 1980					
(1) OLS	-0.020	-0.024	-0.026	-0.019	-0.012	-0.001
	(0.002)	(0.002)	(0.002)	(0.003)	(0.003)	(0.004)
(2) 2SLS	-0.073	-0.026	-0.051	-0.031	-0.054	-0.096
	(0.055)	(0.038)	(0.039)	(0.048)	(0.049)	(0.057)
(2)1st-stage F-statistics	251.80	456.83	471.13	434.68	434.11	397.03
Sample size	117,164	115,706	115,820	108,774	105,827	91,794

The table reports coefficient estimates of Third child

In (2): the instrument for Third child is Same sex

Covariates: Age at first birth, Cohort dummies 1945-1957 (women); 1940-1957 (men), Boy 1st, Boy 2nd

Table A3: Estimates of Labour Force Participation (with Years of Schooling 1990)

Dummy var =1 if	1980	1985	1990	1995	2000	2005
labour earnings > 0	Women 23-35 in 1980					
(1) OLS	-0.047	-0.050	-0.023	-0.023	-0.018	-0.018
	(0.002)	(0.001)	(0.001)	(0.001)	(0.002)	(0.002)
(2) 2SLS	0.032	-0.065	0.006	0.032	0.008	0.026
	(0.048)	(0.025)	(0.018)	(0.025)	(0.028)	(0.036)
(2)1st-stage F-statistics	252.16	408.86	383.74	376.12	373.83	373.09
Sample size	103,966	103,966	103,966	103,966	103,966	103,966
	Men 23-40 in 1980					
(1) OLS	-0.001	-0.007	-0.004	-0.006	-0.006	0.002
	(0.0005)	(0.0006	(0.001)	(0.001)	(0.001)	(0.002)
(2) 2SLS	0.001	-0.013	0.006	-0.016	0.017	0.009
	(0.012)	(0.010	(0.011)	(0.020)	(0.025)	(0.034)
(2)1st-stage F-statistics	261.38	461.75	477.32	458.01	458.15	461.31
Sample size	119,976	119,976	119,976	119,976	119,976	119,976

The table reports coefficient estimates of Third child and Years of Schooling 1990

In (2): the instrument for Third child is Same sex

Covariates: Age at first birth, Cohort dummies 1945-1957 (women); 1940-1957 (men), Boy 1st, Boy 2nd

Table A4 : Estimates of Log Labour Earnings (with Years of Schooling 1990)

Log labour earnings	1980	1985	1990	1995	2000	2005
	Women 23-35 in 1980					
(1) OLS	-0.417	-0.341	-0.194	-0.150	-0.089	-0.056
	(0.008)	(0.006)	(0.004)	(0.005)	(0.005)	(0.006)
(2) 2SLS	-0.256	-0297	-0.067	-0.091	-0.024	-0.046
	(0.183)	(0.096)	(0.069)	(0.081)	(0.080)	(0.097)
(2)1st-stage F-statistics	202.39	367.12	391.99	382.58	363.59	333.78
Sample size	92,640	97,125	100,688	97,815	95,859	89,413
	Men 23-40 in 1980					
(1) OLS	-0.042	-0.055	-0.056	-0.068	-0.053	-0.012
	(0.003)	(0.003)	(0.003)	(0.005)	(0.005)	(0.007)
(2) 2SLS	0.022	0.001	-0.021	0.011	-0.064	-0.084
	(0.072)	(0.056)	(0.054)	(0.079)	(0.083)	(0.105)
(2)1st-stage F-statistics	262.06	451.88	482.04	439.68	438.90	430.95
Sample size	119,236	118,743	118,311	114,324	111,244	100,000

The table reports coefficient estimates of Third child and Years of Schooling 1990

In (2): the instrument for Third child is Same sex

Covariates: Age at first birth, Cohort dummies 1945-1957 (women); 1940-1957 (men), Boy 1st, Boy 2nd

Table A5: Estimates of Labour Force Participation (with Level of Schooling 1990)

Dummy var =1 if	1980	1985	1990	1995	2000	2005
labour earnings > 0	Women 23-35 in 1980					
(1) OLS	-0.047	0.050	-0.023	-0.022	-0.017	-0.018
	(0.002)	(0.001)	(0.001)	(0.001)	(0.002)	(0.002)
(2) 2SLS	0.033	-0.064	0.006	0.031	0.007	0.025
	(0.048)	(0.025)	(0.018)	(0.025)	(0.028)	(0.036)
(2)1st-stage F-statistics	253.46	412.79	387.98	380.27	377.97	377.23
Sample size	103,966	103,966	103,966	103,966	103,966	103,966
	Men 23-40 in 1980					
(1) OLS	-0.001	-0.001	-0.004	-0.007	-0.006	0.002
	(0.0005)	(0.0006)	(0.0006)	(0.001)	(0.001)	(0.002)
(2) 2SLS	0.001	-0.013	0.006	-0.015	0.019	0.011
	(0.012)	(0.010)	(0.011)	(0.020)	(0.025)	(0.035)
(2)1st-stage F-statistics	259.27	459.86	475.49	456.15	456.31	459.49
Sample size	119,976	119,976	119,976	119,976	119,976	119,976

The table reports coefficient estimates of Third child, the instrument for Third child is Same sex in (2)

Covariates include Level of Schooling 1990 and same as in Table 12, the reference is primary education

Table A6: Estimates of Log labour Earnings (with Level of Schooling 1990)

Dummy var =1 if	1980	1985	1990	1995	2000	2005
labour earnings > 0	Women 23-35 in 1980					
(1) OLS	-0.423	-0.347	-0.200	-0.157	-0.095	-0.063
	(0.008)	(0.006)	(0.004)	(0.005)	(0.005)	(0.006)
(2) 2SLS	0.243	-0.290	-0.063	-0.085	-0.019	-0.039
	(0.182)	(0.094)	(0.069)	(0.080)	(0.080)	(0.096)
(2)1st-stage F-statistics	203.74	370.64	395.49	386.46	367.41	169.41
Sample size	92,640	97,125	100,688	97,815	95,859	89,413
	Men 23-40 in 1980					
(1) OLS	-0.045	-0.060	-0.060	-0.075	-0.060	-0.021
	(0.003)	(0.003)	(0.003)	(0.005)	(0.005)	(0.007)
(2) 2SLS	0.017	-0.004	-0.022	0.008	-0.063	-0.089
	(0.072)	(0.056	(0.054)	(0.079)	(0.083)	(0.105)
(2)1st-stage F-statistics	259.91	449.51	480.08	437.49	437.25	429.87
Sample size	119,236	118,743	118,311	114,324	111,244	100,000

The table reports coefficient estimates of Third child, the instrument for Third child is Same sex in (2)

Covariates include Level of Schooling 1990 and same as in Table 12, the reference is primary education