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ABSTRACT

Although theoretical models of labor supply and the family are well developed, there are few credible estimates of key empirical relationships in the work-family nexus. This study uses a new instrumental variable, the sex composition of the first two births in families with at least two children, to estimate the effect of additional children on parents' labor supply. Instrumental variables estimates using the sex mix are substantial but smaller than the corresponding ordinary least squares (OLS) estimates. Moreover, unlike the OLS estimates, the female labor supply effects estimated using sex-mix instruments appear to be absent among more educated women and women with high-wage husbands. We also find that married women who have a third child reduce their labor supply by as much as women in the full sample, while there is no relationship between wives' child-bearing and husbands' labor supply. Finally, we compare these results to estimates produced using twins to generate instruments. Estimates using twins instruments are very close to the estimates generated by sex-mix instruments, once the estimators are corrected for differences in the ages of children whose birth was caused by the instruments. The estimates imply that the labor supply consequences of child-bearing disappear by the time the child is about 13 years old.

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An understanding of the relationship between children and labor supply is important for a number of theoretical and practical reasons. First, economists and demographers have developed a variety of models linking the family and the labor market. Empirical studies of children and labor supply are sometimes seen as tests of these models (e.g., Gronau, 1973; Rosenzweig and Wolpin, 1980a; Schultz, 1990). Second, the link between children and labor supply might partly explain the post-war increase in women's labor force participation rates if having fewer children causes an increase in labor force attachment (Coleman and Pencavel, 1993). Evidence for this includes Goldin (1995, page 29), who has shown that few women in the 1940s and 1950s birth cohorts were able to combine child-rearing with strong labor-force attachment. Other researchers have also drawn a link between fertility-induced withdrawals from the labor force and lower wages of women (e.g., Gronau, 1988; Korenman and Neumark, 1992). So perhaps children keep women from developing their careers.

Any success in disentangling the causal mechanisms linking fertility and labor supply should shed light on other substantive issues as well. For example, reductions in female labor supply could increase the total time parents devote to child care, making at least some children better off (see, e.g., Blau and Grossberg, 1992; Stafford, 1987). Some theories of family behavior also suggest that changes in wives' earnings affect marital stability (Becker, 1985; Becker, Landes and Michael, 1977). Not surprisingly, given the wide and long-standing interest in the relationship between children and labor supply, hundreds of empirical studies report estimates of this relationship. The vast majority of these studies find a negative correlation between fertility (or family size) and female labor supply.[1] As noted in two recent literature surveys, however, the interpretation of these correlations remains unclear. In his assessment of the "economics of the family," Willis (1987, p. 74) writes, "it has proven difficult to find enough well-measured exogenous variables to permit cause and effect relationships to be extracted from correlations among factors such as the delay of marriage, decline of childbearing, growth of divorce, and increased female labor force participation . . ." Browning (1992, p. 1435) expresses similar views: ". . . although we have a number of robust correlations,

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[1]There is relatively little work on the effects of children on husbands' labor supply. See Pencavel (1986, Table 1.17) for a few estimates, which suggest a positive association between fathers' labor supply and the number of children. The relationship between husbands' and wives' labor supply is discussed by, among others, Ashenfelter and Heckman (1974), Heckman and MaCurdy (1980), and Gruber and Cullen (1996).

there are very few credible inferences that can be drawn from them."[2]

Skepticism regarding the causal interpretation of associations between fertility and labor supply arises in part from the fact that there are strong theoretical reasons to believe that fertility and labor supply are jointly determined (see, e.g., Goldin, 1990, p. 125). In fact, this endogeneity is reflected in the academic research agenda. On one hand, papers on labor supply often treat child-status variables as regressors in hours of work equations, while on the other hand, economic demographers and others discuss regressions that are meant to characterize the impact of wages or measures of labor-force attachment on fertility. Since fertility variables cannot be both dependent and exogenous at the same time, it seems unlikely that both types of regressions have a causal interpretation.

This paper focuses on the causal link running from fertility to the work effort of both men and women. Our main contribution is the use of a new instrumental variables (IV) strategy based on the sibling sex mix in families with two or more children. This instrument exploits the widely observed phenomenon of parental preferences for a mixed sibling-sex composition. In particular, parents of same-sex siblings are significantly and substantially more likely to go on to have an additional child.[3] Because sex mix is virtually randomly assigned, a dummy for whether the sex of the second child matches the sex of the first child provides a plausible instrument for further child-bearing among women with at least two children. Moreover, in spite of the fact that the sibling sex mix is obviously a function of the sex of both children, an indicator for having either two boys or two girls is in principle orthogonal to the sex of each sibling. There is therefore little possibility that any secular impact of child-sex on family life contaminates the IV estimates.

We also compare results generated using sex mix as an instrument to results generated using multiple births to construct instruments. Twinning has been used to estimate the causal effects of fertility before, but

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[2]The survey by Nakamura and Nakamura (1992) argues that a search for exogenous variation is so difficult it is not even fruitful (pp. 60-61).

[3]Westoff, Potter, and Sagi (1963) were among the first to report preferences for a mix. In a survey of desired fertility and a follow-up study of actual fertility among couples with two children, they found that parents of two boys or two girls both desired and ultimately had more children than parents of mixed pairs. See Williamson (1976) for an international review. After completing the first draft of this paper in July 1996, we learned of concurrent work using sex-preference instruments to estimate the effect of fertility on female labor supply in the UK (Iacovou, 1996).

never with samples as large or representative as those analyzed here.[4] The juxtaposition of IV estimates based on twinning and the sex mix allows us to compare the effects of children on labor supply when the children are different ages. Combining the two types of instruments, we estimate the time it takes for the labor supply consequences of child-bearing to disappear.

The paper begins with a simple model of fertility, labor supply, and the home production of child care that is used to sketch possible theoretical relationships of interest. Section II discusses the data and the sex-mix instruments' first-stage. Section III presents the main set of empirical results on fertility and labor supply, including an analysis of effects in subgroups defined by husbands' earnings and mothers' schooling. Section IV compares the estimates using sex-mix instruments to estimates based on twins. Section V discusses the empirical findings in light of the theoretical framework and recent trends in female labor force participation. Section VI concludes.

I. Theoretical framework

What are the factors affecting the relationship between children and their parents' labor supply? How are exogenous changes in fertility likely to be reflected in labor supply variables? Does it matter what the source of these changes is? We explore these questions using a framework that incorporates features of the Becker and Lewis (1973) and Becker and Tomes (1976) quantity/quality model into Gronau's (1977) model of home production. The family is treated as an economic unit obtaining utility from leisure and children, with an option to buy or produce an input that increases utility from children. To highlight the key relationships of interest with a minimum of technical detail, we assume additively separable family preferences for leisure, the utility from child services scaled by parental inputs, and a pure child component. Because we are interested in the impact of sex mix and twins-at-second birth, the model describes the choices

[4] Rosenzweig and Wolpin (1980a) used 87 US twin-pairs to estimate labor supply effects, and Rosenzweig and Wolpin (1980b) used 25 twin-pairs from India to estimate the effect of family size on school progress. Neither study presents simple IV estimates. Bronars and Grogger (1994) use twins at first birth from the 1980 Census as instruments to estimate the effect of additional children on *unwed* mother's labor market status. Gangadharan and Rosenbloom (1996) use small samples from the 1980 and 1990 PUMS to estimate the reduced-form effect of twinning on labor supply variables, but they fail to scale the reduced form effects of twinning into effects of childbearing.

facing parents who have already had some children (n_x) but may decide to have additional children (n_c).

The utility function is given by

$$u_\ell(\ell_1, \ell_2) + \beta \ln(n-\gamma) + u_q(nq), \tag{1}$$

where ℓ_1 and ℓ_2 are the fathers' and mothers' leisure time, $n=n_x+n_c$ is the total number of children, and q is a good that increases the utility parents' receive from children. Thus, the sub-utility function for children has the form, $\beta\ln(n-\gamma)+u_q(nq)$, where β and γ are the usual positive Stone-Geary parameters. The commodity q represents what is known as "child quality" in the literature started by Becker and Lewis (1973). Browning (1992) proposes a more neutral terminology, suggesting that the interaction between n and q be viewed simply as a restriction on preferences over two complementary goods. In any case, (1) captures the notion that utility from children is not only a function of numbers, it is also a function of parental inputs. To ensure that second-order conditions hold, we assume that $u_\ell(\ell_1, \ell_2)$ has a negative definite Hessian and that $u_q(nq)$ is increasing and concave.

Fathers' and mothers' time (T) is allocated between work in the market (h_{mj}; j=1,2), child care, i.e., work in the home production of q (h_{hj}; j=1,2), and leisure (ℓ_1, ℓ_2). The production technology for q is given by

$$q = f_1(h_{h1})/n^{\alpha_1} + f_2(h_{h2})/n^{\alpha_2} + q_m;\ 0<\alpha_1 \leq 1,\ 0<\alpha_2 \leq 1 \tag{2}$$

where q_m is purchased per-child inputs and $f_1(h_{h1})$ and $f_2(h_{h2})$ convert parental time spent on child care into an aggregate input. If $\alpha_1=\alpha_2=1$, this aggregate input is simply divided among the children. More realistically, there are likely to be economies of scale in the parental production of q, corresponding to α_1 and α_2 less than 1.

Following Gronau (1977), we assume that q_m can be purchased at fixed prices, p_q, but that $f_1(h_{h1})$ and $f_2(h_{h2})$ exhibit diminishing returns. The model is completed by the family budget constraint, which is

$$p_n n + p_q n q_m = w_1(T-h_{h1}-\ell_1) + w_2(T-h_{h2}-\ell_2) + y, \tag{3}$$

where y is non-labor income and p_n is a fixed per-child cost. This budget constraint captures the Becker and Lewis (1973) idea that the marginal (money) cost of a child depends on parental inputs, q, while the marginal cost of per-child inputs depends on the number of children.

We consider two sources of exogenous variation in the number of children that might generate changes in leisure time and/or work effort. Since $n = n_x + n_c$, where n_x is an exogenously given or predetermined number of children and n_c is chosen by parents, it seems reasonable to view twins as a shock to n_x. Of course, if n_c is unrestricted, any shock to n_x is of no consequence since it can be easily neutralized. But the additional (implicit) restriction $n_c>0$ means that fertility increments induced by twinning cannot always be offset. And as a practical matter, families that experience multiple births do end up having, on average, more children than other families.

The effect of sex preferences on fertility can be modeled in a number of ways. Ben-Porath and Welch (1980) describe the sex mix as something that determines quality in a quantity-quality model. But their model does not put a price on q or allow for home production, so that the sex mix is just a change in utility for inframarginal children based on children's characteristics. The same idea can be captured here by assuming that the sub-utility function corresponding to children distinguishes between marginal and inframarginal children as follows:

$$\beta \ln([1-\theta]n_x+n_c-\gamma) = \beta \ln(n-\gamma^*), \tag{4}$$

where $\gamma^* \equiv \gamma + \theta n_x$.

The parameter θ takes on values between 0 and 1 and discounts inframarginal children if they are same-sex. Thus, the additional fertility that arises as a consequence of sex preferences can be viewed as an attempt to compensate for $[1-\theta]n_x$. In contrast, the additional fertility that arises because of twinning can be viewed as a failure to offset unexpected dn_x.

The fertility and labor supply choices generated by this model can be characterized as follows. Choosing ℓ_1, ℓ_2, n_c, q_m, h_{h1}, and h_{h2} to maximize (1) subject to (2) and (3) determines the relationship between home production and the number of children:

$$f_1'(h_{h1}) = w_1/[p_q n^{1-\alpha_1}] \tag{5}$$

$$f_2'(h_{h2}) = w_2/[p_q n^{1-\alpha_2}]$$

For both partners, time spent in home production is determined solely by real wages (with child care costs as numeraire), marginal productivity at home, and the number of children. If wages are high enough or day care is cheap enough, home time may be zero for either or both partners.

Increasing the number of children generally increases home production, depending on the scale parameters, α_1 and α_2. Differentiating (5), we have

$$dh_{h1} = (\alpha_1 - 1)(w_1/[p_q n^{2-\alpha_1}])(1/f_1'')dn > 0 \tag{6}$$

$$dh_{h2} = (\alpha_2 - 1)(w_2/[p_q n^{2-\alpha_2}])(1/f_2'')dn > 0$$

Only in the case where there are no scale economies should more children leave the time spent in home production unaffected. Note also that, other things equal, higher wages magnify the home production response to changes in the number of children.

Given the identity $T - h_{h1} - \ell_1$, changes in labor supply occur because of changes in home time and leisure time. Equation (6) implies that any shock to n of a given size, whether induced by dn_x or θ, has the same direct effect on home time and hence, labor supply. Moreover, separability between leisure and other goods means that any effect of n on leisure works solely through the marginal utility of income, λ. Therefore, if we are concerned with a marginal-utility-of-income-constant compensated response of the sort commonly considered in the life-cycle labor supply literature (see Browning, Deaton, and Irish, 1985) equation (6) captures the entire labor supply consequences of a change in the number of children.

When λ is not held constant, there are indirect effects on leisure that occur because of the budgetary consequences of any particular shift. From the first order conditions for ℓ_1 and ℓ_2, we have,

$$\begin{matrix} d\ell_1 \\ d\ell_2 \end{matrix} = \begin{bmatrix} u_{\ell 11} & u_{\ell 12} \\ u_{\ell 12} & u_{\ell 22} \end{bmatrix}^{-1} \begin{matrix} w_1 d\lambda \\ w_2 d\lambda. \end{matrix} \tag{7}$$

The exact relationship between $d\lambda$ and $d\ell_j$ clearly depends on both partner's wages and the shape of the sub-utility function for leisure. If husbands' and wives' leisure time are substitutes, then the impact of changes in the marginal utility of income on leisure can operate in different directions for husbands and wives. For example, husbands may take less leisure to offset the wife's decline in earnings from increasing home time and/or leisure time. On the other hand, given the negative-definite Hessian, if $u_{\ell 12}>0$, then we have $d\ell_j/d\lambda < 0$

for j=1,2. In either case, however, the relationship between n and λ depends on the source of variation in n.

To characterize the relationship between n and λ, we can use the first-order conditions for q_{lm} and n_c to write a modified Stone-Geary "demand for children" equation,

$$n = \gamma + \theta n_x + \beta/\{\lambda[p_n-(1-\alpha_1)f_1 n^{-\alpha_1}-(1-\alpha_2)f_2^{-\alpha_2}]\}. \tag{8}$$

Solving for λ, we have

$$\lambda = \beta/[g(n)(n-\gamma^*)] \tag{9}$$

where $g(n)=[p_n-(1-\alpha_1)f_1 n^{-\alpha_1}-(1-\alpha_2)f_2^{-\alpha_2}]$. Differentiating (9) one can show that the sign of $d\lambda/dn$ is theoretically ambiguous, so that the leisure consequences of additional child-bearing are ambiguous even if further restrictions are made on the derivatives of $u_\ell(\ell_1, \ell_2)$. Moreover, in contrast with the home-time relationship expressed in (6), the magnitude of $d\lambda/dn$ depends on whether the source of the change in n is n_x or θ. This dependence arises because the marginal utility consequences of dn_x depend on the value of θ (which discounts n_x) while the marginal utility consequences of $d\theta$ depend on the number of children, n_x.[5]

In summary, this model suggests that the direct labor supply consequences of an exogenous increase in fertility, i.e., those that work by changing the time spent in home production, are similar regardless of source. Labor supply changes that arise because of changes in leisure time also share a common mechanism (changes in the marginal utility of income), but the leisure response is ambiguous and can differ depending on the source. Of course, individuals also have different wages, home production technology, and preferences, all of which can lead to a heterogeneous behavioral response. The model outlined here serves to highlight the individual characteristics that are likely to be the most important determinants of the labor supply response.

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[5]Note that $d\gamma^*/d\theta= n_x$ while $d\gamma^*/dn_x=\theta$.

II. Data, descriptive statistics, and first-stage relationships

A. *Data and descriptive statistics*

The sex-mix IV estimation strategy is implemented using information on labor supply, the sex of mothers' first two children, and an indicator of multiple births in the 1980 and 1990 Census Public Use Micro Data Samples (PUMS). To motivate the empirical work, Table 1 reports labor force participation rates and the probability of additional child-bearing among women aged 21-35 with at least two children in the 1970, 1980, and 1990 PUMS. Data for 1970 are from the 1/100 State file; data for 1980 and 1990 are from the 5 percent samples (Bureau of the Census, 1983 and 1995). The table shows substantial declines in fertility and increases in participation. Statistics for all women aged 21-35 show a similar pattern, as do the statistics for women aged 36-50.

There is no other retrospective fertility information in the PUMS data sets. We therefore matched children to mothers within households in a manner similar to that described in the appendix to Angrist and Evans (1996). Briefly, we attached people in a household labeled as "child" in the primary relationship code to a female householder or the spouse of a male householder. In households with multiple families, we used detailed relationship codes as well as subfamily identifiers to pair children with mothers. We deleted any mother for whom the number of children in the household did not match the reported number of children ever born.[6] Using the sex of the oldest two children, we defined same-sex sibling pairs in both censuses. We defined multiple births in the 1980 Census using the age of children in quarters. There is no information on quarter of birth in the 1990 PUMS, so we were unable to produce an accurate measure of multiple births in 1990.

Because the Census does not track children across households, the sample is limited to mothers aged 21-35 whose oldest child was less than 18 years of age at the time of the Census. Few women younger than 21 have two children, while a child over age 17 is increasingly likely to have moved to a different household. Restricting the women's age group to less than or equal to 35 means the age 18 cutoff for first-born children

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[6]Note also that the sample is restricted to women for whom we observe the age and sex characteristics of her two oldest children and for whom these variables were not allocated by the Census Bureau.

does not generate a highly selected sample. Data from the June 1990 Current Population Survey (CPS) show that among women aged 35 with 2 or more children, at least 93 percent have an oldest child younger than age 18. This fraction falls to 85 percent at age 36 but is equal to 100 percent for women aged 32 or younger. Although women aged 21-35 with at least two children may appear to constitute an unusually young high-fertility group, our tabulations of the June 1990 CPS show that over half of all women aged 28-35 fall into this group. The proportion is lower for women aged 21-27 but still includes at least one quarter of the entire age cohort.

The empirical analysis is conducted on two subsamples from each census data set. The first includes all women with two or more children. The second includes information only on married women and their spouses because this is the sample that the theory is meant to describe. The married sample is also used to explore the impact of children on fathers' labor supply. For the 1980 data, the married sample is restricted to couples who were married at the time of the census, married only once, and married at the time of their first birth. There are 394,835 observations in the full 1980 sample and 254,654 observations in the 1980 married sample (64 percent of the total).[7] Information on the timing of first marriage and the number of marriages is not available in the 1990 PUMS, so that the 1990 married sample includes all women who were married at the time of the Census. The full 1990 sample includes 380,007 women and the married 1990 sample includes 301,588 women (79 percent of the total).

Descriptive statistics and variable definitions for covariates, instruments, and dependent variables are given in Table 2. The covariate of primary interest in our labor supply models is the indicator *More than 2 kids*. The first instrumental variable for *More than 2 kids* is the indicator *Same sex*. The table also shows averages for the two components of *Same sex*, the indicators *Two boys* and *Two girls*. Among all women with two children in 1980, 40.2 percent had a third child. The corresponding figure for 1990 is 37.5. In both samples, just over 50 percent of all two-child families had children of the same sex and just over 51 percent of first births were boys.

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[7]Conditioning the sample on marital status raises the possibility that sample selection bias affects IV estimates in the selected sample. In practice, however, we found no evidence of a relationship between selection status and the instruments.

Labor supply estimates are also computed using multiple births to generate instruments. The mean for our indicator of twin births, which we call *Twins-2,* is 0.0085 in the 1980 full sample and .0083 in the 1980 married sample. For purposes of comparison, we drew a sample of all second births born to women aged 20-35 from the Vital Statistics Natality Data tapes for the years 1971 and 1976 (NCHS, various years). These data sets contain a 50 percent sample of all births in the country and should provide a good alternate estimate of twinning probabilities. The vital statistics data indicate a second-birth twinning probability of 0.0085 for this population. The much higher twinning probabilities for 1990 in Table 1 reflect the fact that no quarter of birth information is available in the 1990 PUMS to identify multiple births. For this reason, we restricted our study of twins to 1980.

Demographic and labor supply variables, described in the lower half of Table 2, include measures of mother's age, age at first birth, years of education, and indicators for race and ethnic background. We also report values for the husbands of women in the married sample. The labor supply variables are based on Census questions concerning work in 1979 or 1989. These variables measure whether respondents *Worked for pay*, their *Weeks worked*, usual *Hours/week*, and annual *Labor income*. The latter three variables are set to zero for those who did not work for pay during the year. The final two variables in the table are measures of *Family income* and, for the married sample, a variable called *Non-mom income* computed as family income minus the wife's labor income.[8] The descriptive statistics show that women's labor force participation rates, weeks and hours worked, and age at first birth increased between 1980 and 1990. Women's real (1995 dollar) earnings increased substantially as well, especially for married women, while real *Non-mom income* declined.

[8]In the few cases where there were negative income values, we set the variables equal to one so that log income is defined in the regressions. Family income and person wage and salary income are topcoded at $75,000 in the 1980 Census. In the 1990 Census, family income is topcoded at $999,999 and person wage and salary income is topcoded at $140,000 with state medians substituted for the topcode.

B. Sex mix and fertility

The phenomenon of parental preferences for a mixed sibling sex composition has been documented in a number of studies. For example, Ben-Porath and Welch (1976) found that in the 1970 Census, 56 percent of families with either two boys or two girls had a third birth, whereas only 51 percent families with one boy and one girl had a third child. Table 3 reports estimates for 1980 similar to these. The first panel looks at sex preferences in families with one or more children by showing the fraction of women with at least one child who had a second child, according to the sex of the first child. The third row of this panel shows the difference by sex. In spite of the fact that attitudinal surveys suggests that many couples would prefer more boys than girls, or prefer their firstborn child to be male (Williamson, 1976), in both samples and both data sets, the fraction of women who had a second child is invariant to the sex of the first child.

The second panel of Table 3 documents the relationship between the fraction of women who have a third child and the sex of the first two children. The first three rows from this section show the sample characteristics for women in the following groups: those with one boy and one girl, those with two girls, and those with two boys. The next two rows report values estimates for women with two children of the same sex and for women with one boy and one girl. The final row reports the differences between the same-sex and mixed-sex group averages.

Both data sets and both samples suggest that women with two children of the same sex are much more likely to have a third child than the mothers of one boy and one girl. For example, in the 1980 all-women sample, only 37.2 percent of women with one boy and one girl have a third child, compared to 43.2 for women with two girls or two boys. The relationship between sex mix and the probability of additional child-bearing is even larger for married women, reaching a precisely estimated 7 percentage point difference for married women in the 1990 Census. This is approximately 21 percent of the rate of additional childbearing among women with one boy and one girl. Finally, we note that the relationship between sex mix and child-bearing is confirmed in data from the fertility supplements to the June 1980, 1985, and 1990 CPS. This is important because, unlike the Census where information about children is partly based on our household match, the June CPS contains detailed fertility histories for each woman including information on the dates of birth and sex of each child.

III. Fertility and labor supply

A. Graphical analysis

Because sibling sex composition is close to being randomly assigned, simple statistical techniques can be used to illustrate how the sex-mix IV strategy identifies the effect of children on parents' labor supply. Consider the bivariate regression model,

$$y_i = \alpha + \beta x_i + \epsilon_i \tag{10}$$

where y_i is a measure of labor supply and x_i is the covariate of interest, *More than 2 kids*. Let z_i denote the binary instrument, *Same sex*. The IV estimate of β in this equation is

$$\beta_{IV} = [(\bar{y}|z_i=1) - (\bar{y}|z_i=0)] / [(\bar{x}|z_i=1) - (\bar{x}|z_i=0)]$$

where $(\bar{y}|z_i=1)$ is the mean of y_i for those observations with $z_i=1$ and other terms are similarly defined. The numerator and denominator capture the reduced-form relationships between y_i and z_i and between x_i and z_i, respectively. The IV method attributes any effect of z_i on y_i to the effect of x_i on y_i.

Although equation (10) is written as a bivariate regression with constant coefficients, Imbens and Angrist (1994) have shown that β_{IV} can be interpreted as a local average treatment effect specific to the instrument, z_i. In this case, β_{IV} estimates the average effect of x_i on y_i for individuals whose value of x_i (the number of children) has been affected by the instrument z_i (sex mix). Similarly, when z_i is the indicator of multiple births, *Twins-2*, the IV estimates reflect the effect of children on labor supply for those who have had more children than they otherwise would have because of twinning.

Figure 1 shows the components of β_{IV} by single year of birth. For example, Figure 1a plots the mean of *More than 2 kids* by *Same sex*. Because few young women have had two kids, the number of women in each age cohort with more than 2 kids increases substantially with age. In all age cohorts, women with two children of the same sex are much more likely to have a third birth. At the bottom of the graph, we report the denominator of the Wald estimate for the entire sample, $[(\bar{x}|z_i=1) - (\bar{x}|z_i=0)]$, which equals 0.06 .

In addition to having more children, women from almost all age cohorts with two children of the same sex have a lower probability of working, work fewer weeks per year, and have lower annual earnings than do women with one boy and one girl. This can be seen in Figures 1b-1d, which plot labor supply outcomes by birth cohort and *Same sex*. The figure captions report the numerator of the associated Wald estimates, $[(\bar{y}|z_i=1) - (\bar{y}|z_i=0)]$, plus the Wald estimate and standard error (labeled b_{iv}). The Wald estimate of the effect of *More than 2 kids* on *Worked for pay* is -0.008/0.06=-0.13, with a standard error of 0.025. The Wald estimates for *Weeks worked* and *Labor income* are -6.4 (1.2) and 2208 (569). The sample mean of *Worked for pay* is 0.57, so that the Wald estimate of -0.13 implies that the presence of a third child reduces the probability a woman works by 23 percent.

Figure 2 shows a similar set of plots using the 1990 data. The relationship between *Same sex* and *More than 2 kids* is larger in 1990 than in 1980, but the reduced form effects of *Same sex* on labor supply variables are somewhat smaller. Both sets of reduced forms therefore imply a slightly smaller IV estimate of the effect of additional childbearing on labor supply and wages.

B. OLS and IV estimates

Linear probability and regression models are used to link labor supply variables for husbands and wives to the endogenous *More than 2 kids* variable and a list of exogenous covariates. The covariates are mother's age and age at first birth, sex of the first-born (*Boy 1st*), and race and Hispanic dummies. The first-stage estimates in Table 4 show that women in 1980 with same sex children are estimated to be 6.1 percentage points more likely to have a third child in a model with covariates. The corresponding estimate for married women is 6.9 percent. As in Table 3, the estimates for 1990 in Table 4 are somewhat larger in both the full and married women samples.

Table 4 also provides some evidence of an association between having a male first-born child and reduced childbearing at higher parities. Note, however, that the effect of *Boy 1st* in the 1980 data is explained entirely by the difference in the effect of *Two boys* and *Two girls* when these regressors are entered separately. In other words, when the effects of sex mix are allowed to differ by sex, there is no relationship between *Boy 1st* and fertility, although the effect of *Same sex* on fertility in 1980 is larger for boys than for girls. The *Boy 1st* effects for 1990 remain significant in all specifications, but they are very small, especially in comparison to the effects of the sex mix.

Next, we use the sex mix to construct instruments for the effect of *More than 2 kids* on measures of employment and earnings in 1980. Table 5 reports a set of OLS estimates and two sets of IV estimates using *Same sex* and the pair of dummies *Two boys* and *Two girls* as instruments. The exogenous regressors are the same as in Table 4 (coefficients not reported). The first three columns show results for the full sample, the next three columns show results for married women, and the last three columns show results for the husbands of married women.

OLS estimates in both the full and married-women sample suggest that the presence of a third child reduces the probability of work by roughly 17 percentage points, weeks worked fall by about 9 per year, hours per week fall by about 7, and family income falls by 13-14 percent. OLS estimates of earnings effects are $3,145 in the married sample and $3,762 in the full sample. Not surprisingly, all of these OLS estimates are very precisely estimated.

In contrast with the results for women, OLS estimates of the effect of *More than 2 kids* on fathers' labor supply are small. Having a third child is estimated to reduce the probability a father worked for pay by less than one percentage point. The impact of a third child on other measures of fathers' labor supply is also small, though precise enough to be significantly different from zero. The estimated effect on annual weeks worked is -.90 and the estimate for hours per week is .16. The effect on husbands' earnings appears substantial (-$1,765) but this amount is still only about 4 percent of the average earnings of men in the sample.

The first set of IV estimates uses *Same sex* as an instrument for *More than 2 kids*. In the full sample, the IV estimates (standard errors) for the dependent variables *Worked for Pay, Weeks worked,* *Hours/week,* and *Labor income* models are -0.12 (0.025), -5.7 (1.1), -4.6 (.95), and -1,961 (544). These results suggest that having a third child causes a 20-30 percent reduction in women's labor supply and earnings. One important finding is that IV estimates using *Same sex* as an instrument are smaller than the corresponding OLS estimates. This is true for the IV labor supply estimates in the married-women's sample as well, although here the gap between IV and OLS is not as large. Overall, however, the OLS estimates appear to exaggerate the causal effect of fertility on female labor supply.

The IV estimates using *Same sex* also differ from most of the IV estimates previously reported in the literature on children and labor supply. In his review article, Browning (1992, p.1469) notes that, "There is one salient difference between studies that take fertility as exogenous and those that take it as endogenous. In many of the latter it is found that fertility either has no effect on labor supply...or it has a positive effect." Browning also points out that it is not clear from these estimates whether children really have no effect on female labor supply, or whether the instruments are too weak or simply poorly chosen. While the IV estimates generated by *Same sex* are smaller than the corresponding OLS estimates, they are still negative, precise, and of a plausible magnitude.

In contrast to the female labor supply estimates, there is little evidence of a relationship between having a third child and family income. Given the strong labor supply effects, the weak impact on family income may seem surprising. There are a few potential explanations for this result. First, the lost income due to a reduction in mothers' labor supply could be made up by other family members. The small and statistically insignificant IV estimates in the *ln(Non-mom income)* equations suggest that this is not the case. The most likely explanation is that the instrument is not powerful enough to detect the family-income consequences of child-bearing. For example, in the married women sample, the third child reduces female earnings by about 21 percent and female labor income is on average 17 percent of total family income. If the third child does not alter the father's labor supply, we would expect an IV coefficient on *More than 2 kids* in the *ln(Family income)* equations of roughly $0.21*.17=0.036$, which is close to the reported estimate of 0.05. But the standard error for this estimate is slightly higher than .05 so that effects this small cannot be precisely measured.

Table 5 also reports estimates of the impact of the third child on fathers' labor supply in the 1980 married sample. While the OLS estimates consistently show a small but significant relationship between fathers' labor supply and additional children, estimates constructed using *Same sex* as an instrument generate no evidence of any effects on the labor supply of men. It is worth noting that the standard errors on IV estimates for fathers' variables *Worked for pay*, *Weeks worked* and *Hours/week* are actually smaller than the corresponding standard errors for women, and they are small enough so that modest positive or negative effects could be detected if they existed.

The labor supply effects estimated using 1990 data are remarkably similar to those estimated for 1980. This can be seen in Table 6, which reports OLS and IV estimates using *Same sex* as an instrument. Some of the estimated effects are slightly smaller in 1990 than in 1980, but these differences are not statistically meaningful. One difference between the 1980 and 1990 results that does seem noteworthy is the larger negative impact of child-bearing on married women's earnings in 1990, perhaps because of an increase in women's wages.

Two boys and *Two Girls*, the components of *Same sex*, can be used as two separate instruments. As we saw in Table 4, mothers of two girls are more likely than mothers of two boys to have a third child. However, the 2SLS estimates in Tables 5 and 6 show that the additional predictive power provided by separating the two components of *Same sex* does not change the coefficient estimates very much or lead to an appreciable increase in precision using either the 1980 or 1990 data. The over-identification test statistic p-value, reported in square brackets in both Tables 5 and Table 6, measures the statistical significance of the difference between IV estimates computed using only *Two boys* as an instrument and IV estimates using only *Two girls* as an instrument. The p-values for the 1990 estimates provide no evidence that it matters which instrument is used. The p-values for some of the 1980 estimates indicate a significant difference in the two sets of estimates. Examining these separate estimates, we found that both the *Two boys* and *Two girls* instruments are always associated with more children and reduced labor supply. The 1980 IV estimates are smaller, however, when *Two girls* is the instrument.

C. Specification issues

The specification issues considered here include the robustness of the results and the validity of the instruments. The robustness of the findings so far is documented in Table 7, which reports IV estimates from models with alternate sets of covariates in the married 1980 sample. Two of the covariates, years of education and fathers' earnings, are potentially endogenous because they may be partly determined by fertility. For this reason they were excluded from the main set of estimates.

The first row of the table reports the same basic IV estimates shown in Table 5 for the married women sample where *Same Sex* is used as the sole instrument for *More than 2 kids*. The remaining estimates are from models that add the following sets of covariates sequentially: i) mother's education, ii) quadratic terms in mom's age, age at first birth, and education iii) linear and quadratic terms in dad's age, dad's age at first birth, and education iv) linear and quadratic terms in dad's labor income, and v) a full set of state dummy variables. The IV estimates are remarkably insensitive to the list of covariates, including the potentially endogenous schooling variables.

The virtual random assignment of *Same sex* makes it very likely that the reduced-form regressions of fertility and labor supply outcomes on the instruments have a causal interpretation. On the other hand, random assignment alone does not guarantee that the <u>only</u> reason the instruments affect labor supply is changing fertility. As with other instruments, there is no direct consistent test for the underlying exclusion restrictions.[9] One possible problem with the exclusion of *Same sex* from labor supply equations is the existence of secular effects of child sex on family life. Reasons for such effects include the possibility that male children reduce the likelihood of divorce by increasing the father's commitment to the family (see, e.g., Morgan, Lye, and Condran, 1988). It is also possible that children's sex is connected with family background variables that are associated with individual or ethnic differences in the timing and frequency of sexual

[9]The distinction between randomization and exclusion restrictions is discussed by Angrist, Imbens, and Rubin (1996). Exclusion restrictions do have testable implications for certain inequalities; see Imbens and Rubin (1994) for details.

intercourse (Weinberg, Gladen, and Wilcox, 1994; Guttentag and Secord, 1983). Last, there could be secular impacts of sex mix generated by the fact that boys are more likely than girls to have disabilities.

Bias from pure child-sex effects of this kind is impossible if the probability of having a male child is exactly .5, because in that case *Same sex* is orthogonal to the sex of each child. But any deviation from this proportion leads to correlation between *Same sex* and the sex of each child. To see this formally, let s_1 and s_2 be indicators for the sex of first and second-born children. Note that

$$\textit{Same sex} = s_1 s_2 + (1-s_1)(1-s_2).$$

The covariance of s_j and *Same sex* is $E[s_1 s_2]-E[s_j]E[\textit{Same sex}]$, which equals $E[s_j](E[s_j]-E[\textit{Same sex}])$ assuming that child sex is independent and identically distributed (i.i.d.) over children. This can be further simplified to show that the regression of *Same sex* on s_j produces a slope coefficient equal to $2E[s_j]-1$, which is zero if $E[s_j]=1/2$.[10] Since the probability of giving birth to a male child is .51, there should be a slight positive association between *Same sex* and the sex of each child, and that is the case for our data.

To assess the extent of bias from pure child-sex effects, Table 8 contrasts results generated by using *Same sex* as an instrument in models with no controls for the sex of either child to results generated using models with additive controls for the sex of both first and second-born children. As it turns out, the IV estimates are almost completely invariant to the inclusion of regressors that control for the sex of each child. Even more remarkably, there is not much evidence of an association between having male children and labor supply. The only significant own-sex coefficients are for the effect of *Boy 2nd* in the 1980 sample, but these effects are small. *Boy 2nd* effects on earnings, not shown in the table, are not significant in either data set.

Finally, we note that Butcher and Case (1994) have argued that sibling sex composition affects girls' schooling because girls with brothers may be more likely to be raised like boys (and hence to get more schooling or at least be more likely to go to college). But Kaestner (1995) fails to replicate their findings using data for a more recent cohort. Also, the fact that we find similar parental labor supply and fertility

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[10]Proof: Assuming child sex is i.i.d., we have $E[s_1]=E[s_2]$ and $E[s_1 s_2]=E[s_j]^2$. Therefore, $Cov(\textit{Same sex}, s_j) = E[s_j](E[s_j]-E[\textit{Same sex}])$. Some manipulation gives $E[s_j]-E[\textit{Same sex}]=(1-E[s_j])(2E[s_j]-1)$. Since the variance of s_j is $E[s_j](1-E[s_j])$, the regression coefficient is $(2E[s_j]-1)$.

effects in families with both boy pairs and girl pairs seems to weigh against theories based on differences in the way parents treat male and female children.

D. Additional results

The theoretical discussion focuses attention on how the effect of children on labor supply might vary with the wages or earnings potential of either husbands or wives. The first panel of Table 9 reports OLS and IV estimates of the effect of *More than 2 kids* for married women, conditional on the position of their husbands in the husbands' earnings distribution. The first column shows the first-stage relationship between *More than 2 kids* and *Same sex*, interacted with dummies that indicate whether husbands earnings are in the upper third, middle third, or lower third of the earnings distribution. These estimates show that the effect of *Same sex* on fertility is increasing in husbands earnings. For women with high-wage husbands, however, the labor supply effects are close to zero and are not significantly different from zero. With the exception of weeks worked in 1980, the effects are also larger for middle-income families than for low-income families. Note that average participation rates do not decline enough with husbands earnings to account for the lack of effects among women with high wage husbands. It is also worth noting that, in contrast to the IV results, the OLS estimates show large and significant labor supply effects at all levels of husbands' earnings.

It is not possible to analyze the labor supply effects conditional on women's wages because wages are unobserved for women who don't work. But we can condition on schooling, which is an important predictor of individual earnings potential. This is done in panel b of Table 9 for married women with less than a high school education (18 and 16 percent of the 1980 and 1990 samples respectively), high school graduates (49-38 percent of the samples), and more than a high school education (33-45 percent of the samples). The reduced forms show a strong association between *Same sex* and fertility in each schooling group, although the effect is about 1 percentage point smaller for the mothers in the highest education category. The IV estimates suggest that women with relatively low levels of schooling experience the largest effects of children on labor supply. There is no statistically significant association between additional childbearing and labor supply for women with more than a high school education. As with the estimates that conditioned on husbands' labor supply, the schooling-group variation in the IV estimates differ from the OLS estimates (not reported in the table), which show similar effects at all schooling levels.

Because mothers' education and husbands' wages are correlated, it is not clear whether a set of estimates that condition on husbands' earnings and a set of estimates that condition on mothers' education are capturing distinct phenomena. We therefore present estimates by mothers' education-group in a sample restricted to women whose husbands have earnings in the middle third of the earnings distribution. As in panel b, these estimates suggest that child-bearing has no effect on the labor supply of the most educated women, with monotonically declining effects as education rises. This finding is even more remarkable when viewed in light of the fact that participation rates increase with mothers' schooling. Finally, we estimated effects at each third of the husbands earnings distribution in a sample of mothers that includes high school graduates only. These results are similar to those reported in panel a of Table 9.

IV. Comparison with estimates using multiple births

An alternative and equally plausible instrument for the effect of child-bearing on labor supply is multiple births. Figure 3a shows the relationship between a multiple second birth and the probability of having more than two children in the 1980 Census. This probability is identically equal to one for women who experience a multiple birth. For other women, it is less than one but greater than zero and increasing with age. Overall, the effect of multiple second births on the probability of having more than two children is about .6.

Women who experience a multiple second birth are less likely to be working, work fewer weeks per year, and have lower earnings than other women with 2 or more children who do not experience a multiple second birth. This can be seen in Figures 3b, 3c, and 3d, which plot reduced-form relationships and report the implied Wald estimates of labor supply and earnings effects. Wald estimates (standard errors) for effects on *Worked for pay*, *Weeks worked*, and *Labor income* are -0.08 (0.014), -3.3 (0.6), and -947 (313). These effects are all statistically significant, but not as large as the estimates generated by the *Same sex* instrument.

A full set of IV estimates using *Same sex* and *Twins-2* are compared in Table 10. As in Table 4, the models used to produce these estimates include a variety of exogenous covariates to control for mothers' age, race, age at first birth, and the sex of the first-born child. Additional covariates included in these models are the ages of the first and second child in quarters. The estimates of female labor supply effects produced using *Twins-2* are consistently smaller than the corresponding estimates using *Same sex*. Although the contrast between *Same sex* and *Twins-2* coefficient estimates is not large enough to be statistically significant for many of the individual coefficients, the set of estimates generates a pattern that strongly suggests these two shocks have different effects.

A likely explanation for the smaller *Twins-2* effects is that, conditional on the age of the second child, a third child who is born as a consequence of twinning is necessarily older than a third child who is born for other reasons. The reason for this is that third children who are born as twins are exactly the same age as second children, while some time must pass for a non-twin third child to be born. In the 1980 sample, for example, the average age of third children who are twins is 6.4 years while the average age of other third children is 5 years. Regression-adjusting for the covariates used to construct the estimates in Table 10, the age gap between twins and other third children grows to almost three years. This difference in ages has implications for labor supply estimates if the effect of children on labor supply is larger when the children are younger.

We use the following model to check whether differences in the *Same sex* and *Twins-2* IV estimates can be explained by differences in the ages of third children. The equation of interest is,

$$y_i = \alpha' w_i + \beta_i x_i + \epsilon_i, \tag{11}$$

where w_i is a vector of exogenous covariates and x_i is the endogenous regressor *More than 2 kids*. The coefficient β_i is an individually-varying causal effect that depends on the age of the third child. In particular, we assume

$$\beta_i = \beta_0 + \beta_1 a_{3i} \tag{12}$$

where a_{3i} is equal to the age of the mothers' third child for women who have a third child and equal to zero otherwise. Combining (11) and (12) generates the estimating equation,

$$y_i = \alpha' w_i + \beta_0 x_i + \beta_1 (a_{3i} x_i) + \epsilon_i. \tag{13}$$

Assuming that differences in a_{3i} are the only reason why the *Same sex* and *Twins-2* instruments generate different estimates, we can use both instruments to estimate the coefficients on the two endogenous regressors in (13), x_i and $a_{3i}x_i$.

IV estimates of β_0 and β_1 are reported in Table 11 for the full 1980 sample, where a_{3i} was measured in quarters for the purposes of estimation. All of the estimates of β_0 are negative and all of the estimates of β_1 are positive, suggesting that the negative impact of childbearing declines as the third child ages. The table also reports estimates of the value of a_{3i} at which β_i=0. This is $a^* \equiv -\beta_0/\beta_1$. Estimates of a^* are 12.6 years for effects on *Worked for pay*, 12.4 years for effects on *Weeks worked*, 14.4 years for effects on *Hours/week*, and 12.9 years for effects on *Labor income*. It seems like a good approximation to say that the effect of the third child on labor supply decays to zero around the time the child turns 13. Assuming this is true, we can obtain more precise estimates of β_0 (and β_1) by imposing the restriction $\beta_1 = -\beta_0/13$. The F-tests for this restriction suggest that it fits models for each dependent variable well. Once restricted, the estimates of β_1 and β_0 are also very precise.[11]

To illustrate how this model reconciles the *Same sex* and *Twins-2* estimates, note that if $\beta_1 = -\beta_0/a^*$, we have

$$y_i = \alpha' w_i + \beta_0 (1 - a_{3i}/a^*) x_i + \epsilon_i. \tag{14}$$

Choosing a value for a^*, we can use the *Same sex* and *Twins-2* instruments to construct separate estimates of β_0 in (14) by treating $(1 - a_{3i}/a^*)x_i$ as the single endogenous regressor. The results when a^*=13, reported in panels b and c of Table 11, show that the *Same sex* and *Twins-2* instruments generate very similar estimates of β_0 in (14) for all dependent variables. This suggests that the model of the effect of child-bearing embodied

[11]The linear model is obviously an approximation since it implies that the effect of childbearing on labor supply becomes positive once the third child is older than a^*. Only about 2 percent of third children are older than age 13 in our data because the oldest mother is age 35.

in (12), combined with the restriction that β_i decays to zero at age 13, does a good job of reconciling the *Same sex* and *Twins-2* estimates.[12]

V. Implications

A. Fertility and the rise in female labor supply

At the turn of the century, less than 20 percent of all workers were women. Today, women make up almost half the work force (Goldin, 1990). A number of labor economists and economic historians have attempted to decompose the rise in the female labor force participation rates into components attributable to demand and supply shifts. For example, Mincer (1962) found that 90 percent of the rise in postwar labor force participation of married women can be attributed to an increase in demand. Smith and Ward (1984) also found that demand characteristics can explain a majority of the increase in total hours worked by all women in between 1850 and 1980. In contrast, Goldin (1990) argues that shifts in supply explain about half of the change in female labor force participation between 1960 and 1980.

Declining fertility represents a potentially important supply shifter that might account for some of the increase in female labor attachment. How much of the trend in labor force attachment in the population we have studied can be accounted for by reduced childbearing beyond the second child? In Table 1, we showed that the probability of having more than two children for women aged 21-35 with at least two children fell by 18 percentage points between 1970 and 1990, a drop of about one-third. At the same time, labor force participation rates rose by 21 percentage points, a 48 percent increase. Similar statistics for other groups reported in Table 1 show that our sample is not unusual in experiencing these trends.

Using the *Same sex* IV estimate of the impact of *More than 2 kids* on *Worked for pay* from table 4 (-.121), declining fertility can account for an employment increase equal to .18*.121, which is about 2 percentage points. This calculation suggests that even though childbearing clearly affects labor supply,

[12]If we set $a^*=-\beta_0/\beta_1$ using the coefficient estimates from each equation, then the *Same sex* and *Twins-2* estimates of (14) for any equation will necessarily be identical. The point of estimation with a^* fixed at 13 is to show that one extra free parameter (β_1), combined with a single plausible value for a^*, reconciles all of the *Same sex* and *Twins-2* estimates.

declining fertility can explain only a small fraction of the overall rise in female labor force participation.

B. Theories of children and economic behavior

The model outlined here predicts that additional child-bearing will increase the time spent in home production for both partners, although reductions in leisure time could offset some of the negative labor supply consequences of these changes. The IV estimates generated using *Same sex* and *Twins-2* confirm a negative effect of children on mothers' labor supply. The theoretical model also predicts that fertility increments attributable to sex preferences and twinning should have similar labor supply consequences. This seems to be the case once the estimates are adjusted for age differences among third children.

The lack of a relationship between the number of children and fathers' labor supply seems harder to reconcile with the theoretical framework, which predicts effects of childbearing on the home production and leisure time of both fathers and mothers. Finally, the model predicts that the impact of children on labor supply should be increasing in the wages of both partners, while the IV estimates suggest the opposite is true: more educated women and/or women with the most highly paid husbands exhibit the smallest labor-supply response to the third child. In fact, many of the estimates for women with more than a high school education or whose husbands have earnings in the top third of the earnings distribution are virtually zero.

VI. Conclusions

Economic models of household behavior generate a rich variety of predictions and theoretical relationships, few of which have been confronted with credible empirical evidence. The evidence reported here is unique in that it derives from plausibly exogenous natural experiments in family size. However, the empirical results probably raise as many questions as they answer.

IV estimates that exploit the sex mix and twinning both confirm the OLS estimates showing that children lead to a reduction in female labor supply, although the OLS estimates may exaggerate the causal effect of children. This is probably not too surprising, at least not to the mothers of small children. What is surprising is that the effects of children on labor supply are absent among certain groups. Equally important is the finding that husbands do not change their labor market behavior in response to a change in family size. Families absorb the cost of extra child care for the third child either through a reduction in the wife's earnings, with no offsetting increase in husbands' earnings, or by purchasing child care services from non-family providers. These findings suggests that in spite of the increase in womens' wages and labor force participation rates, traditional family roles may still be more important for the intra-family allocation of time than economic considerations.

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Figure 1: Means of Fertility and Labor Market Outcomes,
Women Aged 21-35 With 2 or More Children,
By Sex of First Two Children,
1980 PUMS

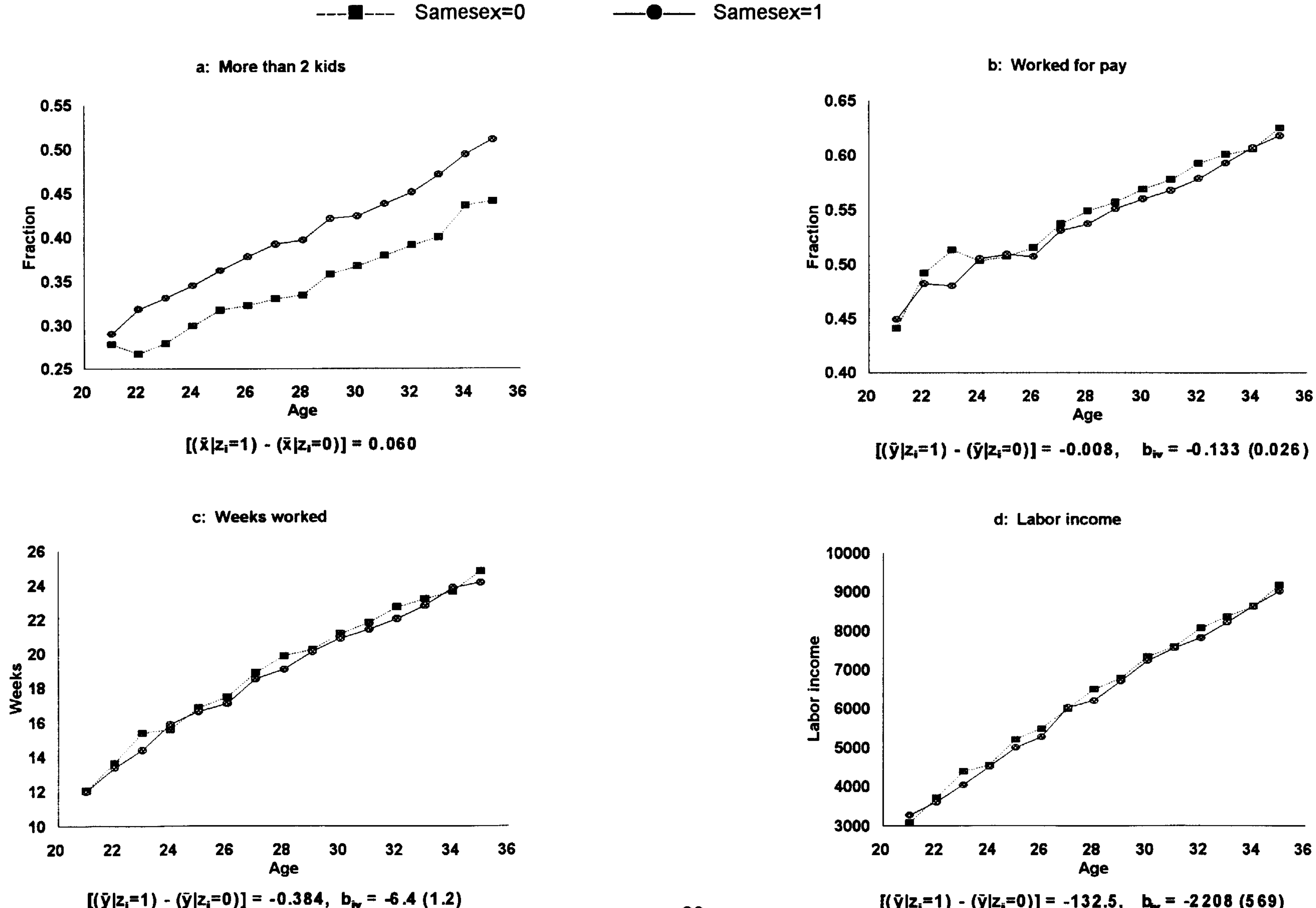


Figure 2: Means of Fertility and Labor Market Outcomes,
Women Aged 21-35 With 2 or More Children,
By Sex of First Two Children,
1990 PUMS

---■--- Samesex=0 ——●—— Samesex=1

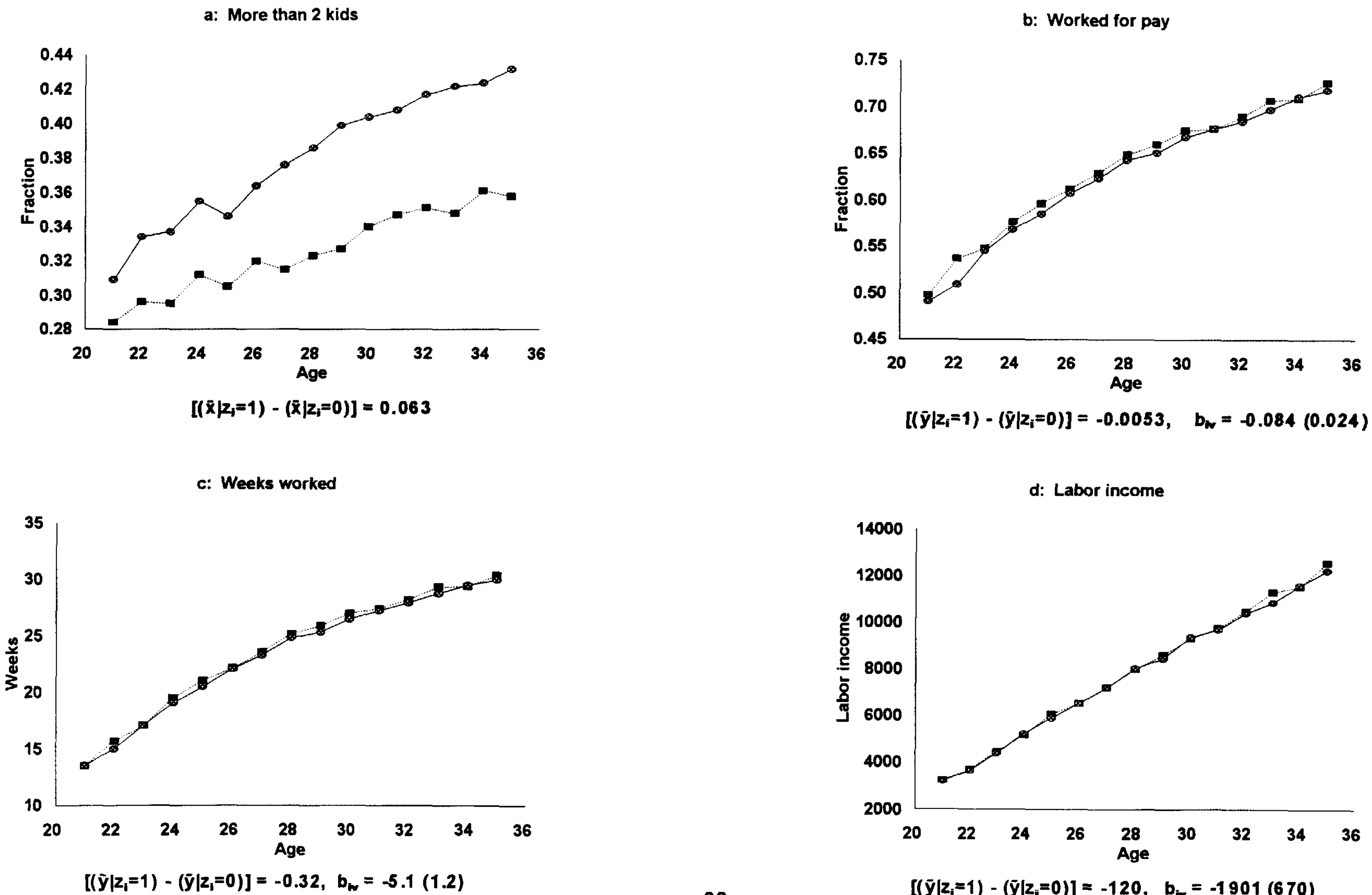
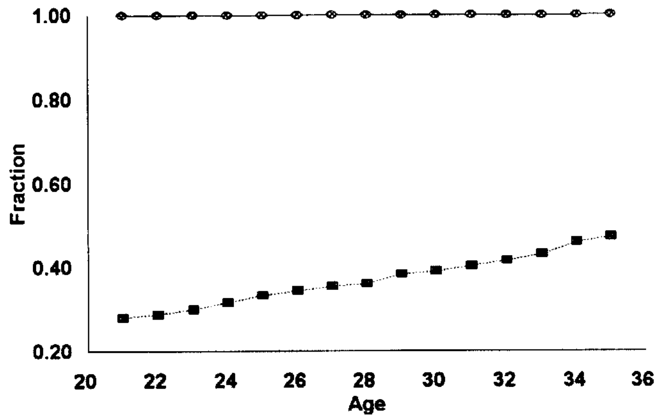


Figure 3: Means of Fertility and Labor Market Outcomes,
Women Aged 21-35 With 2 or More Children,
By Twin Status of Second Birth
1980 PUMS

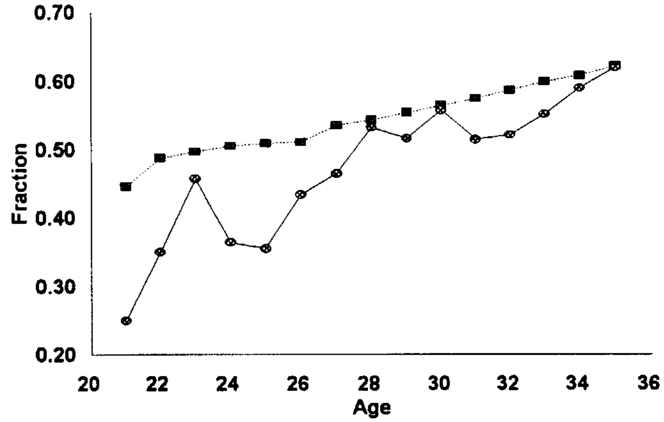
---■--- Twin -- 2=0 —●— Twin -- 2=1

a: More than 2 kids



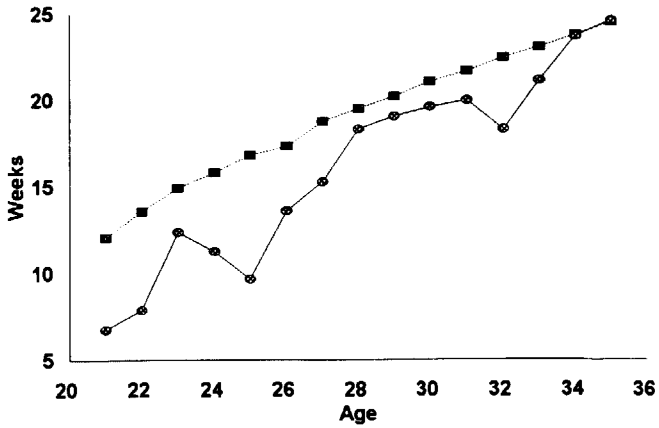
$[(\bar{x}|z_i=1) - (\bar{x}|z_i=0)] = 0.603$

b: Worked for pay



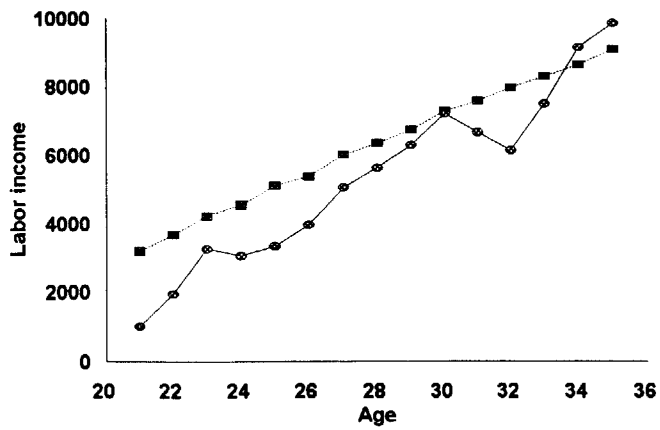
$[(\bar{y}|z_i=1) - (\bar{y}|z_i=0)] = -0.046$, $b_{iv} = -0.076$ (0.014)

c: Weeks worked



$[(\bar{y}|z_i=1) - (\bar{y}|z_i=0)] = -1.98$, $b_{iv} = -3.3$ (0.6)

d: Labor income



$[(\bar{y}|z_i=1) - (\bar{y}|z_i=0)] = -570.7$, $b_{iv} = -946.5$ (313)

Table 1
Fertility and Labor Supply Measures,
Women Aged 21-50,
1970, 1980, and 1990 PUMS

Sample	1970 PUMS	1980 PUMS	1990 PUMS
Women aged 21-35			
Mean children even born	1.78	1.27	1.18
% with 2 or more kids	52.1%	40.4%	37.6%
% worked last year	60.0%	73.4%	79.3%
Observations	203,918	1,326,631	1,478,546
Women aged 36-50			
Mean children even born	2.85	2.86	2.15
% with 2 or more kids	74.3%	78.5%	68.9%
% worked last year	57.3%	66.7%	78.5%
Observations	181,502	852,204	1,253,095
Women, aged 21-35 with 2 or more children			
% women with more than 2 kids	55.6%	39.9%	39.1%
% worked last year	44.8%	58.0%	66.6%
Observations	106,239	535,587	577,397
Married women, aged 21-35 with 2 or more children			
% women with more than 2 kids	54.9%	39.0%	37.5%
% worked last year	41.8%	55.8%	67.5%
Observations	91,286	436,483	439,408

The 1970 PUMS data is from the 1/100 State file, from the 5% sample. The married samples include women married at the time of the Census.

Table 2
Descriptive Statistics,
Women Aged 21-35 With 2 or More Children,
1980 and 1990 PUMS

Variable	Means and (Standard Deviations) 1980 PUMS All women	1980 PUMS Married couples Wives	1980 PUMS Married couples Husbands	1990 PUMS All women	1990 PUMS Married couples Wives	1990 PUMS Married couples Husbands
Children ever born	2.55	2.51		2.50	2.48	
	(0.81)	(0.77)		(0.76)	(0.74)	
More than 2 kids	0.402	0.381		0.375	0.367	
(=1 if mother had more than 2 children, =0 otherwise).	(0.490)	(0.486)		(0.484)	(0.482)	
Two boys	0.264	0.266		0.263	0.265	
(=1 if first two children were boys).	(0.441)	(0.442)		(0.441)	(0.441)	
Two girls	0.242	0.239		0.241	0.239	
(=1 if first two children were girls).	(0.428)	(0.427)		(0.428)	(0.426)	
Same sex	0.506	0.506		0.505	0.504	
(=1 if first two children were the same sex).	(0.500)	(0.500)		(0.500)	(0.500)	
Boy 1st	0.511	0.514		0.512	0.514	
(=1 if first born was a boy).	(0.500)	(0.500)		(0.500)	(0.500)	
Twins - 2	0.0085	0.0083		0.012	0.011	
(=1 if second birth was a twin).	(0.0920)	(0.0908)		(0.108)	(0.105)	
Age	30.1	30.3	33.0	30.4	30.7	33.4
	(3.5)	(3.4)	(4.6)	(3.5)	(3.3)	(4.8)
Age at first birth	20.1	20.8	24.0	21.8	22.3	25.1
(Parent's age in years when first child was born).	(2.9)	(2.9)	(4.0)	(3.5)	(3.5)	(4.7)
Worked for pay	0.565	0.528	0.977	0.662	0.667	0.968
(=1 if worked for pay in year prior to census).	(0.496)	(0.499)	(0.150)	(0.473)	(0.471)	(0.175)
Weeks worked	20.8	19.0	48.0	26.2	26.4	47.1
(Weeks worked in year prior to census).	(22.3)	(21.9)	(10.5)	(22.9)	(22.9)	(12.0)
Hours worked/week	18.8	16.7	43.5	22.5	22.1	44.0
(Average hours worked per week).	(18.9)	(18.3)	(12.3)	(19.1)	(18.9)	(13.3)
Labor income	7,160	6,250	38,915	9,550	9,616	36,623
(Labor earnings in year prior to census, in 1995 dollars).	(10,804)	(10,210)	(25,012)	(13,071)	(13,238)	(30,283)
Family income	42,339	47,641		42,558	49,196	
(Family income in year prior to census, in 1995 dollars).	(26,561)	(25,819)		(34,692)	(34,740)	
Non-mom income		41,391			39,580	
(Family income - Mom's labor income, in 1995 dollars).		(24,582)			(31,891)	
Number of observations	394,835	254,654	254,654	380,007	301,588	301,588

Women whose second child is less than a year old are deleted from all samples. In the 1980 PUMS, the married women sample refers to women who were married at the time of their first birth, married at the time of the survey, and married only once. In the 1990 PUMS, the married women are those married at the time of the census. Sample weights are used in all calculations based on the 1990 PUMS.

Table 3
Fraction of Families Who Had Another Child,
By Parity and Sex of Children,
Women Aged 21-35, 1980 and 1990 PUMS

	All women				Married women			
	1980 PUMS (649,887 observations)		1990 PUMS (627,362 observations)		1980 PUMS (410,333 observations)		1990 PUMS (477,798 observations)	
Sex of first child, families with one or more children	(1) Fraction of sample	(2) Fraction who had another child (std. error)	(3) Fraction of sample	(4) Fraction who had another child (std. error)	(1) Fraction of sample	(2) Fraction who had another child (std. error)	(3) Fraction of sample	(4) Fraction who had another child (std. error)
(1) one girl	0.488	0.694 (0.001)	0.489	0.665 (0.001)	0.485	0.720 (0.001)	0.487	0.698 (0.001)
(2) one boy	0.512	0.694 (0.001)	0.511	0.667 (0.001)	0.515	0.720 (0.001)	0.513	0.699 (0.001)
difference (2) - (1)		0.000 (0.001)		0.002 (0.001)		0.000 (0.001)		0.001 (0.001)

Table 3 (continued)

Sex of first two children, families with two or more children		All women				Married women			
		1980 PUMS (394,835 observations)		1990 PUMS (380,007 observations)		1980 PUMS (254,654 observations)		1990 PUMS (301,588 observations)	
		(1) Fraction of sample	(2) Fraction who had another child (std. error)	(3) Fraction of sample	(4) Fraction who had another child (std. error)	(1) Fraction of sample)	(2) Fraction who had another child (std. error)	(3) Fraction of sample	(4) Fraction who had another child (std. error)
	one boy, one girl	0.494	0.372 (0.001)	0.495	0.344 (0.001)	0.494	0.346 (0.001)	0.497	0.331 (0.001)
	two girls	0.242	0.441 (0.002)	0.241	0.412 (0.002)	0.239	0.425 (0.002)	0.239	0.408 (0.002)
	two boys	0.264	0.423 (0.002)	0.264	0.401 (0.002)	0.266	0.404 (0.002)	0.264	0.396 (0.002)
(1)	one boy, one girl	0.494	0.372 (0.001)	0.495	0.344 (0.001)	0.494	0.346 (0.001)	0.497	0.331 (0.001)
(2)	both same sex	0.506	0.432 (0.001)	0.505	0.407 (0.001)	0.506	0.414 (0.001)	0.503	0.401 (0.001)
	difference (2) - (1)		0.060 (0.002)		0.063 (0.002)		0.068 (0.002)		0.070 (0.002)

Samples are defined in the footnotes to Table 2.

Table 4
OLS Estimates of *More than 2 kids* Equations,
Women Aged 21-35 With 2 or More Children ,
1980 and 1990 PUMS

Parameter Estimates (Standard Errors)

	1980 PUMS						1990 PUMS					
	All women			Married women			All women			Married women		
Independent Variable	(1)	(2)	(3)	(1)	(2)	(3)	(1)	(2)	(3)	(1)	(2)	(3)
With other covariates	no	yes	yes	no	yes	yes	no	yes	yes	no	yes	yes
Age		0.0302 (0.0002)	0.0302 (0.0002)		0.0302 (0.0003)	0.0302 (0.0003)		0.0247 (0.0002)	0.0247 (0.0002)		0.0274 (0.0003)	0.0274 (0.0003)
Age at first birth		-0.0451 (0.0003)	-0.0451 (0.0003)		-0.0439 (0.0003)	-0.0438 (0.0003)		-0.0386 (0.0002)	-0.0386 (0.0002)		-0.0390 (0.0003)	-0.0390 (0.0003)
Boy1st		-0.0081 (0.0015)	0.0002 (0.0021)		-0.0113 (0.0018)	-0.0014 (0.0027)		-0.0081 (0.0015)	-0.0083 (0.0022)		-0.0097 (0.0017)	-0.0086 (0.0024)
Same sex	0.0600 (0.0016)	0.0614 (0.0015)		0.0675 (0.0019)	0.0691 (0.0018)		0.0628 (0.0016)	0.0623 (0.0015)		0.0702 (0.0018)	0.0703 (0.0017)	
Two boys			0.0533 (0.0021)			0.0596 (0.0026)			0.0624 (0.0021)			0.0692 (0.0023)
Two girls			0.0698 (0.0021)			0.0792 (0.0027)			0.0621 (0.0022)			0.0714 (0.0024)
R^2	0.004	0.084	0.084	0.005	0.078	0.078	0.004	0.082	0.082	0.005	0.082	0.082

Samples are defined in the footnotes to Table 2. Other covariates in the models include indicators for *Black, Hispanic,* and *Other race.*

Table 5
OLS and 2SLS Estimates of Labor Supply Models,
Women Aged 21-35 With 2 or More Children,
1980 PUMS

Parameter Estimates (Standard Errors) on *More than 2 kids* variable, [p-value, test of over identifying restrictions]

	All women			Married women			Husbands of married women		
Model	(1)	(2)	(3)	(1)	(2)	(3)	(1)	(2)	(3)
Estimation method	OLS	2SLS	2SLS	OLS	2SLS	2SLS	OLS	2SLS	2SLS
Instrument for *More than 2 kids*		*Same sex*	*Two boys, Two girls*		*Same sex*	*Two boys, Two girls*		*Same sex*	*Two boys, Two girls*
Dependent variable:									
Worked for pay	-0.176 (0.002)	-0.121 (0.025)	-0.113 (0.025) [0.015]	-0.167 (0.002)	-0.122 (0.028)	-0.112 (0.028) [0.014]	-0.008 (0.001)	0.003 (0.009)	0.001 (0.008) [0.011]
Weeks worked	-8.98 (0.07)	-5.68 (1.11)	-5.32 (1.10) [0.019]	-8.02 (0.09)	-5.45 (1.21)	-5.14 (1.20) [0.073]	-0.90 (0.04)	0.60 (0.60)	0.41 (0.59) [0.026]
Hours/week	-6.64 (0.06)	-4.61 (0.95)	-4.35 (0.94) [0.032]	-5.96 (0.08)	-4.87 (1.02)	-4.59 (1.01) [0.047]	0.16 (0.05)	0.53 (0.70)	0.45 (0.69) [0.424]
Labor income	-3762.0 (35.5)	-1960.5 (544.1)	-1852.6 (539.4) [0.136]	-3145.7 (42.0)	-1338.0 (571.1)	-1308.4 (565.5) [0.713]	-1764.8 (103.7)	-1316.1 (1405.3)	-1460.1 (1391.0) [0.475]
ln(Family income)	-0.135 (0.004)	-0.030 (0.065)	-0.040 (0.064) [0.266]	-0.138 (0.004)	-0.050 (0.056)	-0.054 (0.056) [0.661]			
ln(Non-mom income)				-0.060 (0.005)	0.026 (0.067)	-0.013 (0.067) [0.249]			

Samples are defined in the footnotes to Table 2. Other covariates in the models include *Age, Age at first birth, Boy 1st,* plus indicators for *Black, Hispanic,* and *Other race.*

Table 6
OLS and 2SLS Estimates of Labor Supply Models,
Women Aged 21-35 With 2 or More Children,
1990 PUMS

Parameter Estimates (Standard Errors) on *More than 2 kids* variable, [p-value, test of over identifying restrictions]

	All women			Married women			Husbands of married women		
Model	(1)	(2)	(3)	(1)	(2)	(3)	(1)	(2)	(3)
Estimation method	OLS	2SLS	2SLS	OLS	2SLS	2SLS	OLS	2SLS	2SLS
Instrument for *More than 2 kids*		*Same sex*	*Two boys, Two girls*		*Same sex*	*Two boys, Two girls*		*Same sex*	*Two boys, Two girls*
Dependent variable:									
Worked for pay	-0.155 (0.002)	-0.092 (0.024)	-0.092 (0.024) [0.743]	-0.147 (0.002)	-0.104 (0.024)	-0.104 (0.024) [0.576]	-0.012 (0.001)	0.017 (0.009)	0.017 (0.009) [0.989]
Weeks worked	-8.71 (0.08)	-5.64 (1.16)	-5.64 (1.16) [0.391]	-8.25 (0.09)	-5.74 (1.15)	-5.76 (1.15) [0.670]	-1.03 (0.05)	1.01 (0.63)	1.05 (0.63) [0.708]
Hours/week	-6.80 (0.07)	-4.10 (0.98)	-4.10 (0.98) [0.489]	-6.39 (0.07)	-3.95 (0.96)	-3.95 (0.96) [0.665]	-0.06 (0.05)	0.85 (0.69)	0.83 (0.69) [0.180]
Labor income	-3984.4 (44.2)	-2096.6 (663.8)	-2096.2 (663.8) [0.830]	-3753.9 (50.8)	-2454.9 (669.7)	-2456.3 (669.7) [0.893]	929.7 (114.9)	1364.4 (1536.1)	1354.8 (1535.9) [0.711]
ln(Family income)	-0.120 (0.005)	-0.122 (0.071)	-0.122 (0.071) [0.270]	-0.103 (0.004)	-0.054 (0.051)	-0.054 (0.051) [0.878]			
ln(Non-mom income)				-0.004 (0.005)	0.019 (0.068)	0.020 (0.068) [0.452]			

Samples are defined in the footnotes to Table 2. Other covariates in the models are listed in the footnotes to Table 5.

Table 7
2SLS Estimates of Female Labor Supply Models,
Married Women Aged 21-35 With 2 or More Children,
1980 PUMS

Parameter Estimates (Standard Errors) on *More than 2 kids* Variable

	Dependent Variable					
Model	*Worked for pay*	*Weeks worked*	*Hours /week*	*Labor income*	*ln(Family income)*	*ln(Non-mom income)*
Basic model, Table 4	-0.122 (0.028)	-5.45 (1.21)	-4.87 (1.02)	-1338.0 (571.1)	-0.050 (0.056)	0.026 (0.067)
Previous model plus mom's years of education	-0.120 (0.028)	-5.38 (1.21)	-4.83 (1.02)	-1295.9 (566.0)	-0.045 (0.055)	0.030 (0.067)
Previous model plus quadratic terms in mom's age, age at first birth, and years of education	-0.122 (0.028)	-5.37 (1.21)	-4.88 (1.02)	-1344.4 (564.0)	-0.043 (0.055)	0.033 (0.067)
Previous model plus linear and quadratic terms in dad's age, age at first birth, and years of education.	-0.120 (0.028)	-5.26 (1.20)	-4.78 (1.01)	-1290.1 (563.7)	-0.048 (0.055)	0.027 (0.067)
Previous model plus linear and quadratic terms in dad's labor income	-0.124 (0.028)	-5.48 (1.20)	-4.99 (1.00)	-1351.7 (562.2)	-0.004 (0.044)	0.083 (0.053)
Previous model plus state effects	-0.124 (0.027)	-5.51 (1.18)	-5.09 (0.99)	-1390.7 (557.5)	-0.006 (0.045)	0.083 (0.053)

Samples are defined in the footnotes to Table 2. Other covariates in the models are listed in the footnotes to Table 5.

Table 8
2SLS Estimates of Labor Supply Models,
Women Aged 21-35 With 2 or More Children,
1980 and 1990 PUMS

Parameter Estimates (Standard Errors)

	1980 PUMS						1990 PUMS					
Variable	*Worked for pay*		*Weeks worked*		*Hours/week*		*Worked for pay*		*Weeks worked*		*Hours/week*	
Boy 1st		-0.0000 (0.0015)		-0.087 (0.069)		-0.032 (0.059)		-0.0019 (0.0015)		-0.074 (0.072)		-0.125 (0.061)
Boy 2nd		-0.0038 (0.0015)		-0.161 (0.069)		-0.126 (0.059)		-0.0005 (0.0015)		0.062 (0.072)		-0.042 (0.061)
More kids	-0.121 (0.025)	-0.119 (0.025)	-5.71 (1.12)	-5.62 (1.11)	-4.63 (0.95)	-4.57 (0.95)	-0.093 (0.024)	-0.092 (0.024)	-5.66 (1.16)	-5.66 (1.15)	-4.14 (0.98)	-4.08 (0.98)

Other covariates in the models include *Age*, *Age at first birth*, plus indicators for *Black*, *Hispanic*, and *Other race*.

Table 9
2SLS Estimates of Labor Supply Models,
Married Aged Women 21-35 With 2 or More Children,
1980 PUMS

Parameter Estimates (Standard Errors)

	1980 PUMS					1990 PUMS				
	More than 2 kids	*Worked for pay*		*Weeks/year*		*More than 2 kids*	*Worked for pay*		*Weeks/year*	
Sample/Variables	First-stage	Mean of dependent variable	2SLS	Mean of dependent variable	2SLS	First-stage	Mean of dependent variable	2SLS	Mean of dependent variable	2SLS
a: Results by *Dad's labor earnings*. *Same sex* or *More than 2 kids* x										
x bottom third of Dad's labor earnings	0.056 (0.003)	0.570	-0.120 (0.060)	21.1	-7.43 (2.60)	0.064 (0.003)	0.669	-0.125 (0.045)	26.3	-5.98 (2.18)
x middle third of Dad's labor earnings	0.072 (0.003)	0.569	-0.187 (0.047)	20.8	-7.16 (2.04)	0.075 (0.003)	0.728	-0.147 (0.039)	29.8	-8.05 (1.87)
x top third of Dad's labor earnings	0.079 (0.003)	0.448	-0.079 (0.042)	15.2	-3.14 (1.83)	0.071 (0.003)	0.610	-0.029 (0.040)	23.6	-2.59 (1.93)
b: Results by *Mom's education*. *Same sex* or *More than 2 kids* x										
Mom < high school graduate	0.071 (0.004)	0.468	-0.119 (0.064)	16.1	-6.91 (2.80)	0.069 (0.004)	0.532	-0.254 (0.061)	19.2	-12.8 (2.91)
Mom high school graduate	0.073 (0.003)	0.524	-0.148 (0.039)	19.2	-6.47 (1.65)	0.078 (0.003)	0.662	-0.094 (0.035)	26.3	-5.47 (1.67)
Mom > high school graduate	0.062 (0.003)	0.567	-0.083 (0.054)	20.4	-2.92 (2.33)	0.064 (0.002)	0.719	-0.057 (0.038)	29.1	-3.52 (1.84)
c: Results by *Mom's eduation* for women with *Middle third of Dad's labor earnings*: *Same sex* or *More than 2 kids* x										
Mom < high school graduate	0.079 (0.008)	0.481	-0.276 (0.108)	16.7	-10.2 (4.80)	0.072 (0.008)	0.579	-0.283 (0.098)	21.7	-15.6 (4.90)
Mom high school graduate	0.076 (0.004)	0.551	-0.192 (0.060)	20.3	-7.91 (2.65)	0.081 (0.004)	0.707	-0.197 (0.051)	28.8	-8.96 (2.58)
Mom > high school graduate	0.062 (0.006)	0.640	-0.126 (0.097)	23.7	-4.02 (4.30)	0.071 (0.043)	0.795	-0.066 (0.059)	33.3	-5.60 (2.96)

Samples are defined in the footnotes to Table 2. Other covariates in the models include those listed in the footnotes to Table 5 plus, when appropriate. indicators of dad's labor earnings or dom's education.

Table 10
2SLS Estimates of Labor Supply Models,
Women Aged 21-35 With 2 or More Children,
1980 PUMS

Parameter Estimates (Standard Errors) on *More than 2 kids* variable,

	All women		Married women		Husbands	
Model	(1)	(2)	(1)	(2)	(1)	(2)
Instrument for *More than 2 kids*	*Same sex*	*Twins--2*	*Same sex*	*Twins -- 2*	*Same sex*	*Twins -- 2*
Dependent variable:						
Worked for pay	-0.125	-0.079	-0.124	-0.086	0.004	-0.001
	(0.026)	(0.013)	(0.028)	(0.017)	(0.009)	(0.005)
Weeks worked	-5.84	-3.63	-5.51	-4.21	0.65	-0.33
	(1.15)	(0.60)	(1.23)	(0.72)	(0.61)	(0.36)
Hours/week	-4.78	-3.33	-4.95	-3.49	0.55	-0.47
	(0.98)	(0.51)	(1.03)	(0.61)	(0.71)	(0.42)
Labor income	- 1962.1	-1258.6	- 323.6	-1466.7	-1258.8	664.7
	(562.8)	(292.6)	(580.9)	(339.8)	(1428.8)	(840.3)
ln(Family income)	-0.020	-0.071	-0.048	-0.024		
	(0.067)	(0.035)	(0.057)	(0.033)		
ln(Non-mom income)			0.028	0.053		
			(0.068)	(0.040)		

Samples are defined in the footnotes to Table 2. Other covariates in the models are listed in the footnotes to Table 4.

Table 11
2SLS Estimates of Two-Parameter Labor Supply Models,
Women 21-35 With 2 or More Children,
1980 PUMS

Parameter Estimates (Standard Errors)

	Worked for pay		*Weeks/year*		*Hours/week*		*Labor income*	
Variable	Not restricted	Restricted	Not restricted	Restricted	Not restricted	Restricted	Not restricted	Restricted
a: Instruments: *Sames sex* and *Twins -- 2*								
β_0	-0.188	-0.182	-8.84	-8.44	-6.73	-7.40	-1397	-1396
	(0.066)	(0.024)	(2.91)	(1.09)	(2.48)	(0.91)	(679)	(249)
β_2	0.015	0.014	0.711	0.649	0.466	0.568	108.2	105
	(0.010)	(0.002)	(0.429)	(0.082)	(0.366)	(0.070)	(49.9)	(14.2)
a^*	12.6	13	12.4	13	14.4	13	12.9	
	(3.98)		(3.60)		(6.26)		(5.83)	
F: $-\beta_1=\beta_0/a^*$		0.011		0.022		0.083		0.011
(p-value)		(0.92)		(0.88)		(0.77)		(0.98)
b: Instrument: *Same sex* (restricted: $\beta_0 = -\beta_1\, a^*$)								
β_0		-0.185		-8.64		-7.07		-1383
		(0.039)		(1.72)		(1.47)		(400)
c: Instrument: *Twins -- 2* (restricted: $\beta_0 = -\beta_1\, a^*$)								
β_0		-0.180		-8.31		-7.61		-1371
		(0.031)		(1.37)		(1.17)		(319)

Samples are defined in the footnotes to Table 2. Other covariates in the models are listed in the footnotes to Table 4.