FEMALE LABOR SUPPLY AND FERTILITY.
CAUSAL EVIDENCE FOR LATIN AMERICA*

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ABSTRACT

This paper studies the causal relationship between fertility and female labor supply using census data from 14 Latin American countries and the U.S. over three decades (1980, 1990 and 2000). Parental preferences for a gender-balanced family is exploited as a source of exogenous variation. Although OLS estimates suggest a negative relationship, instrumental variables fails to identify a causal effect in most countries. When considering a pool of married women, a negative causal effect is found. In any case, despite having a highly accurate first-stage, the analysis of the quality of the instrument reveals a weak explanatory power of sibling sex composition on fertility.

Keywords: Causality, Female Labor Supply, Fertility, Latin America.

RESUMEN

Este artículo estudia la relación causal entre fertilidad y oferta de trabajo femenino usando datos censales de 14 países de Latinoamérica y EE.UU en un período de tres décadas (1980s, 1990s y 2000s). La preferencia de los padres por una estructura familiar balanceada (en términos de género) se utiliza como fuente de variación exógena. Aunque las estimaciones de MCO sugieren una relación negativa, el método de variables instrumentales no identifica un efecto para la mayor parte de los países. Considerando las mujeres casadas se obtiene un efecto negativo. El análisis de la calidad del instrumento revela un débil poder explicativo del género sobre la fertilidad.

*Palabras clave: Causalidad, Oferta de Trabajo Femenina, Fertilidad, América Latina.*

I. Introduction

The increase in the participation of women in the labor market has been one of the most dynamic labor milestones worldwide during the last century. A lot of theoretical and empirical studies have attempted to account for the possible explanations for this increase (Killingsworth and Heckman, 1986; Chioda, 2011). Many of them focused their arguments on the determinants of the demand side, while others did so for the supply side. In particular, a stream of these studies focused on the relationship between the biological phenomenon of the conception of o spring (fertility) and the economic phenomenon of working women, finding a negative and robust correlation between these variables in all of them.

Using data from World Development Indicators (WDI) for several Latin American countries in the period 1980-2009, Figure Nr. 1 shows the evolution of the female labor participation rate (ratio of women working or seeking work in relation to the working age population) and the fertility rate (births per woman). On average, female labor participation has increased monotonically (30%), while fertility has decreased monotonically

Figure Nr. 1: Labor Force Participation and Fertility (Latin America)
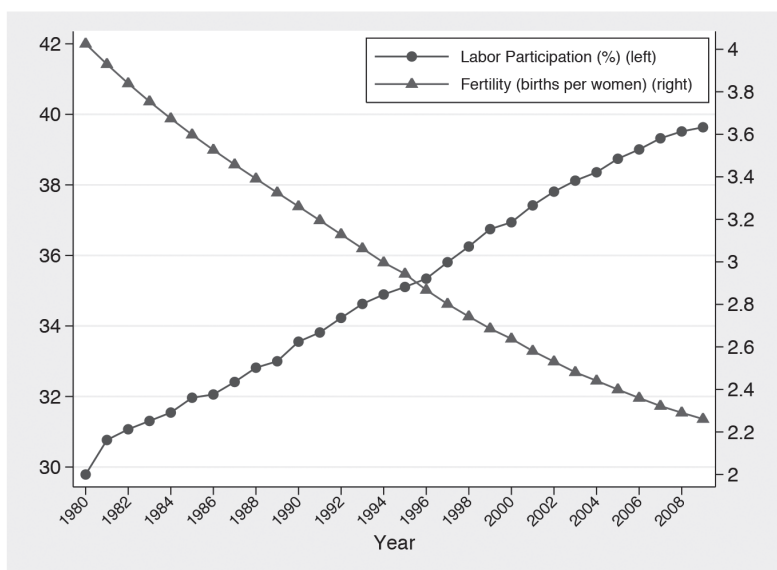
(44%) over the period. This stylized fact is present in each country, as can be seen in Figure Nr. 2.[1]

The main problem that arises from these simple negative correlations lies in the simultaneity between fertility and female labor supply, which prevents the interpretation of this relationship as a causal effect. Moreover, the observed negative correlations between fertility and labor supply could be spurious.

Figure Nr. 2: Labor Force Participation and Fertility (by countries)
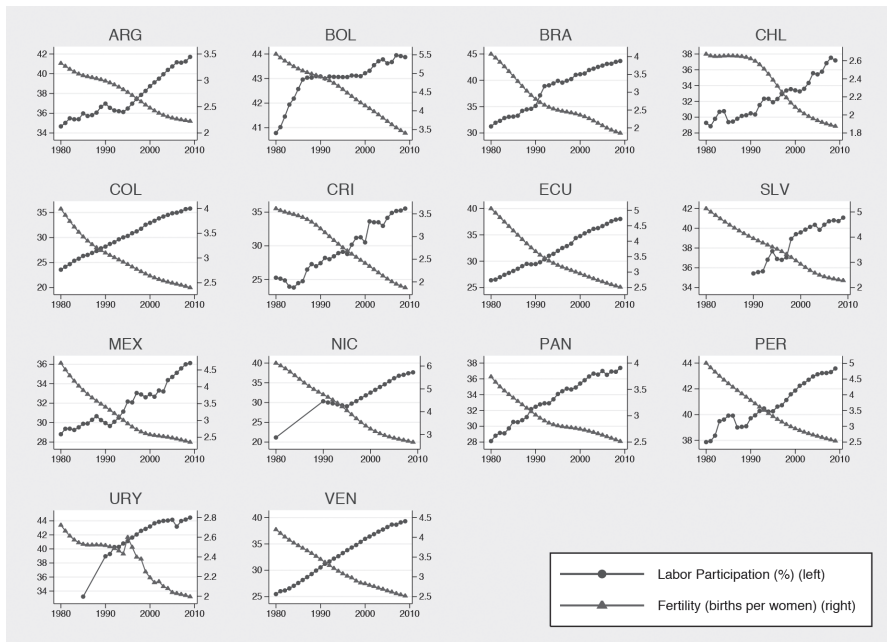
Source: own estimates based on WDI-World Bank. Countries from left to right are: Argentina, Bolivia, Brazil, Chile, Colombia, Costa Rica, Ecuador, El Salvador, Mexico, Nicaragua, Panama, Peru, Uruguay and Venezuela. Left axis: \Female labor participation (% of total labor force)"; Right axis: \Fertility rate (births per woman)".

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1. Guinnane (2011) studies the historical transition of European countries and the United States from high fertility to low fertility between the nineteenth and twentieth centuries. Before the transition, women conceived up to eight children on average and the elasticity of fertility with respect to income was positive. Currently, many women choose not to have children, and the elasticity of fertility with respect to income is zero or even negative.

On the basis of these arguments, Angrist and Evans (1998) (henceforth AE) estimated a negative causal effect of fertility on female labor supply for the U.S. exploiting a source of exogenous variability in family size: the parental preferences for a mixed sibling sex composition (Williamson, 1983). This stylized fact has been documented in numerous studies and indicates that parents of same-sex siblings are significantly more likely to have an additional child.[2] Since the sex mix is virtually randomly assigned, an indicator variable for whether the sex of the second child matches the sex of the first child provides a plausible instrument for further childbearing among women with at least two children, and in this way it is possible to measure the effect of moving from the second to the third child on labor supply.

Many other studies have tried to extend AE's work to other countries using the same source of exogenous variation in fertility. Iacovou (2001) and Van der Stoep (2008) used the same-sex instrument as in AE for the United Kingdom and South Africa, respectively, and in both cases, the IV estimates were not statistically significant. Daouli, Demoussis and Giannakopoulos (2009) did the same for Greece, finding a negative causal effect on the probability of working at the 10% level in 1991, which disappears in 2001. In Sweden, Hirvonen (2009) found a strong negative effect on women's earnings and a mild effect on labor force participation. Cools (2012) obtained a similar result for Norway, although the effect on labor participation is not precise enough to be statistically different from zero. Cruces and Galiani (2007) found a negative causal effect for Argentina and Mexico. Finally, Chun and Oh (2002) for South Korea, and Agüero and Marks (2008) for a pool of 6 Latin American countries used variations of AE's empirical strategy.[3] In the first case, the authors found a strong negative causal effect on labor supply, and in the second case the estimates

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2. *For instance, Ben-Porath and Welch (1976) found for the 1970 census of the United States that 56% of the families whose first two children were of the same sex had a third child, while 51% with a boy and a girl had a third child.*

3. *Chun and Oh (2002) exploit South Korean households' preferences towards male children because of their superior labor market performance with respect to females. They instrument fertility with the sex of the first child under the assumption that if it is a female, their parents will try to conceive another one. In Agüero and Marks (2008) the exogenous source of variation in family size is based on infertility shocks as a random event.*

were imprecise. Taking all these studies into account, it follows that causal evidence between fertility and labor supply is far from being conclusive.

So far, the study of Cruces and Galiani (2007) (henceforth CG) is the only causal evidence that uses AE's identification strategy for Latin American countries. The authors point out that, compared to the U.S. women in Latin America are characterized by having more children, lower education levels and fewer formal facilities for childcare. Likewise, households in Latin America have faced structural changes in recent decades, which have affected women's labor decisions and the allocation of their resources within the household. First, the rise in female labor participation meant a new source of household income. Second, investment in education grew steadily with important consequences not only for women's earning potential but also for their identity and aspirations. Third, there has been an extended effort in the region to reduce poverty through policies that directly or indirectly favored women's access to income and economic assets, such as microcredit programs and conditional cash transfer (CCT) programs (Chioda, 2011). For all these reasons, it is interesting to extend the analysis to a broader group of developing countries.

Lastly, analyzing the relationship between fertility and labor supply could be of political interest since both variables are associated with poverty and well-being. For example, if fertility actually had a negative effect on female labor supply, a system of subsidies for childcare could relax the temporal restriction of mothers, fostering their reinsertion into the labor market, and providing an extra income source for the family.

In this sense, the main purpose of the paper is to proceed in this research line, trying to determine whether the negative causal effect of fertility on women's labor participation found in the U.S., Argentina and Mexico can be extended to other Latin American countries.[4]

Throughout the work, I use census data from 14 Latin American countries and the U.S. (as a benchmark for a developed country) for the 1980s, 1990s and 2000s, and the relationship between the variables of interest is

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4. As mentioned by Angrist (2004), the external validity of IV estimates is ultimately established less by new econometric methods than by replication in new data sets.

estimated by Ordinary Least Squares (OLS) and Two-Stage Least Squares (2SLS).

Even though OLS estimates suggest a negative and statistically significant relationship between fertility and mothers' labor supply in each Latin American country, 2SLS estimates fail to identify a causal effect in most of them. Namely, the average effect of moving from a two-child family to a larger one is statistically zero for those women whose fertility decisions are changed by the instrument (compliers). Considering a sample of married women for a pool of countries over the span of three decades, a negative and statistically significant causal effect is found. In any case, despite having very precise first-stage estimations (and evidence in favor of the exclusion restriction), an analysis of the quality of the instrument reveals a weak explanatory power of sibling sex composition on fertility. The problem of weak instruments entails a huge efficiency loss in the second-stage, with large standard errors that make the interpretation of IV estimates meaningless.

The remainder of the paper is structured as follows: Section II sets out a simple model of female labor supply and introduces the empirical strategy. Section III provides a description of the data sets used for estimation as well as some summary statistics. Later, section IV discusses the internal validity of the strategy and section V presents and discusses the results of my analysis. Section VI presents conclusions.

II. Conceptual and empirical framework
II.1. Theoretical Framework
The relationship between female labor supply and the number of children can be represented by an adapted version of the stylized static model of Browning (1992). The woman's utility function can be defined as $U = u(C, \theta, h)$, where C denotes the consumption of the mother and children with price p_C, θ is the time devoted to leisure, and h is the number of children with a cost per child given by p_h. The function is assumed to be increasing in all its arguments. The woman divides her total time T between work at home $l_f = g(h)$ (housework and childcare), θ leisure and work in the market l_m for which she receives a wage w. Besides, there is a fixed income I from

the household. The logic of the model is that, while children provide utility to their parents, they also enter the household budget constraint since they involve considerable costs, both in terms of goods (e.g., food, clothing and school materials) and time devoted to childcare. Each woman solves the following utility maximization problem:

$$\max_{C,\theta,h} U = u(C,\theta,h) \text{ s.t.} \qquad (1)$$

$$I + wl_m = p_C C + p_h h \quad \textit{(budget constraint)} \qquad (2)$$

$$T = l_f + l_m + \theta \quad \textit{(time constraint)} \qquad (3)$$

The two constraints can be summarized as $I+wT=(w\theta+p_C C)+(p_h h+wl_f)$, which describes the allocation of the household's full income between the woman and the child. An explicit utility function and the first and second order conditions of the optimization problem define the demand for children and the women's labor supply.

Even though this simple setting, as it stands, is too general to derive explicit solutions, it still captures the essence of the theoretical relationship between fertility and female labor supply: the utility function, the budget and time constraints imply a trade-off between "pure" utility from children, labor income and the needs of children (time and goods).[5] Besides, the model can be used to illustrate the underlying endogeneity that arises in the empirical estimation of labor supply models.

This work seeks to identify the direct effect of the children h on the labor supply of women, represented by $l^*_m = T - \theta^* - l^*_f$. Following Browning (1992), the model from equations (1)-(3) results in a conditional labor supply (either in terms of hour -intensive margin- or as a binary participation indicator -extensive margin-) defined as $Y= f(K,D)$, where K is a vector that contains the variables in the model and some exogenous characteristics, and D is a measure of fertility (such as the number of children h, or an indicator of more than h children in a sample of women with h or

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5. *Dynamic models often predict a negative causal effect of fertility on short-run labor supply through the time needs of children in the time constraint.*

more children). The parameter of interest is the labor supply response to changes in fertility, f_D. However, this parameter is difficult to recover by simple statistical methods. For instance, ignoring the effects of fertility on all other variables included in K and considering the potential effects of fertility on wages, results in:

$$\frac{\partial Y}{\partial D} = \frac{\partial w}{\partial D} f_w + f_D \qquad (4)$$

Childbearing might have an effect on wages ($\partial w/\partial D \neq 0$), for example because of the foregone appreciation in the woman's "stock of experience" during maternity leave. Moreover, since wages are determined by ability and motivation or ambition, which are unobservable, they may be correlated with fertility decisions through childbearing and leisure preferences in the utility function U. Taking into account all the variables of the model would add partial derivatives of the components of K with respect to D. This discussion suggests that a fertility indicator D would be endogenous in a labor supply model. An additional factor is that unobserved factors might be driving both decisions.

In this sense, Willis (1987) suggests that a solution to this endogeneity problem is to find a variable Z that induces variation in fertility but does not directly affect labor supply decisions, which allows the derivation of a reduced form relationship between fertility and labor supply. Continuing the example presented in equation (4), if Z is not related to the factors that account for $\partial w/\partial D$, then:

$$\frac{\partial Y}{\partial Z} = \frac{\partial w}{\partial Z} f_w + \frac{\partial D}{\partial Z} f_D \quad \Rightarrow \quad f_D = \frac{\partial Y}{\partial Z} \Big/ \frac{\partial D}{\partial Z} \qquad (5)$$

since the exogeneity of Z with respect to w implies that $\partial w/\partial Z = 0$. In this way, the parameter of interest, i.e. the response of labor supply to changes in fertility, is identified.

II.2. Empirical Strategy
This subsection presents the empirical strategy adopted throughout the work to identify the direct effect of fertility on female labor supply. Based

on equation (4) of the theoretical framework, a first attempt at estimating the mentioned effect consists in comparing the average occupational status of women, by running an OLS regression of Y on D. However, this simple comparison is unlikely to identify any meaningful causal effect due to the presence of selection bias.[6]

To solve this endogeneity problem I rely on sex mix as a natural experiment. This strategy, first proposed by Angrist and Evans (1998), relies on the sex of women's first two children as an instrument for fertility, and can be justified as follows. The sex composition of children affects fertility through parental sex preferences. That is, parents whose first two children have the same sex, exhibit a higher probability of having another child to attain their desired composition (Williamson, 1983). Since the gender of children is random and making the identifying assumption that it affects labor supply only through its effect on fertility, the *Same-sex* indicator can be used as an instrument for fertility and thus a causal effect can be identified for the subpopulation of compliers.[7] Bearing this in mind, I estimate the following equation:

$$Y_i = \alpha' X_i + \alpha_1 s_{1i} + \alpha_2 s_{2i} + \beta D_i + \varepsilon_i \qquad (6)$$

where Y is a measure of labor supply, D is the endogenous fertility measure, X includes plausibly exogenous characteristics like the age of the woman and her age at first birth, and s_1 and s_2 are indicators for the sex of the first two children.[8] This model is estimated by 2SLS with a first-stage regression of the form:

$$D_i = \delta' X_i + \delta_1 s_{1i} + \delta_2 s_{2i} + \gamma Z_i + v_i \qquad (7)$$

6. *For instance, childbearing decisions take into account expected potential outcomes, career plans, comparative advantages, preferences and the division of labor within the household. Then, it could happen that women lacking opportunities for childcare arrangements, with stronger preferences for children, or who forecast relatively poor labor market outcomes (such as a low quality job, or low wages) self-select into treatment (have children) and decide to work at home. On the other hand, women who expect good labor market outcomes probably self-select into lower fertility (non treatment).*

7. *Compliers are women who would have had an additional child if their first two children were of the same sex, but would not have had it if the first two were of different sex.*

8. *AE argue that Same-sex can be written as a function of s_1 and s_2, and therefore it is potentially correlated with the sex of either child. Accordingly it is important to control for any secular additive effect of child sex.*

Since the 2SLS framework allows for more than one instrument, the estimations are also carried out by decomposing the *Same-sex* indicator into *Two-boys*= $s_1 s_2$ and *Two-girls*= $(1-s_1)(1-s_2)$.[9] In this case, the first-stage model is:

$$D_i = \delta' X_i + \delta_1 s_{1i} + \gamma_0 z_{1i} + \gamma_1 Z_{2i} + v_i \qquad (8)$$

The parameters in the first-stage are γ, γ_0 and γ_1. The parameter of interest in the second-stage is β: the average effect of D_i on Y_i for those women whose fertility status is changed by the instrument.

III. Data and summary statistics

The microdata employed in the present study come from the Integrated Public Use Microdata Series International (IPUMS-International), a project dedicated to collecting, harmonizing and distributing census data from around the world. I analyze the data from 14 Latin American countries (Argentina, Bolivia, Brazil, Chile, Colombia, Costa Rica, Ecuador, El Salvador, Mexico, Nicaragua, Panama, Peru, Uruguay and Venezuela) and the U.S. for the 1980s, 1990s and 2000s.[10] In order to carry out the estimations, further adjustments were necessary. Women with at least two children were selected from the total sample, and their characteristics were linked to those of their children. As in AE and CG, the samples were limited to women aged between 21 and 35 years, whose eldest child was not older than 18 years and whose second child was at least 1 year old at the time of the census.[11]

Regarding the variables used in the study, the outcome of interest for the estimations (*Worked for pay*) is an indicator equal to 1 if the woman worked in the reference week (typically the week prior to the census) and is not a family worker without remuneration and 0 otherwise. The reason

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9. When using these two instruments, it is not possible to control for the sex of the first two children because of perfect multicollinearity. As in AE and CG, the results control for the sex of the first child.

10. The data correspond to 10% nationally representative samples with the exception of Brazil and the U.S. where the samples are approximately 5%. A list of countries and years used are reported in Table A1 of the online appendix at http://economics:dtortarolo:com:ar/.

11. Since I use the same data source and make the same adjustments as in AE and CG, the results can be compared with that of them. More details about the adjustments of the data can be found in the online appendix.

for using this variable is that it is available for all countries and periods considered, and it is the same variable used by AE and CG. The fertility variable (*More than 2 children*) is an indicator defined as 1 for women with three or more children, and 0 otherwise. This endogenous indicator is instrumented by *Same-sex*, *Two-boys* and *Two-girls* which are equal to 1 if the first two children were the same sex, two males or two females respectively, and 0 otherwise in all cases.

Before restricting the sample to women with two or more children, it is important to analyze the evolution of fertility and the female labor participation rate using the raw data from IPUMS. The left panel of Figure Nr. 3 (see page 38) summarizes the average information for Latin America and the right panel shows a scatter plot of all the countries and years considered. The pattern obtained matches the WDI data discussed above, where, on average, the participation of women in the labor market increased and the fertility decreased over the last three decades.

Regarding the descriptive statistics in the subsample of mothers, the average labor force participation in Latin America, as captured by the *Worked for pay* indicator, was 20% in 1980, 25% in 1990 and 32% in 2000, with variability between countries that decreases over time. These numbers are lower than those of the U.S. (46% in 1980, 55% in 1990 and 58% in 2000). The average number of children in Latin America was 3.3, 3.1 and 2.8 for 1980, 1990 and 2000, respectively, and higher than the average number of children in the U.S. (approximately 2.6 in the three censuses). In the case of the *More than 2 children* indicator, 63% of mothers with at least two children had a third child in 1980, 58% did so in 1990 and 50% in 2000. In the U.S. these percentages are barely 41%. Finally, it is worth noting that on average Latin American mothers are less educated than those of the U.S., although there is much variability across countries.[12]

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12. *The tables with the information for each country can be found in the online appendix.*

**Figure Nr. 3: Female labor participation and number of children
(Latin America)**
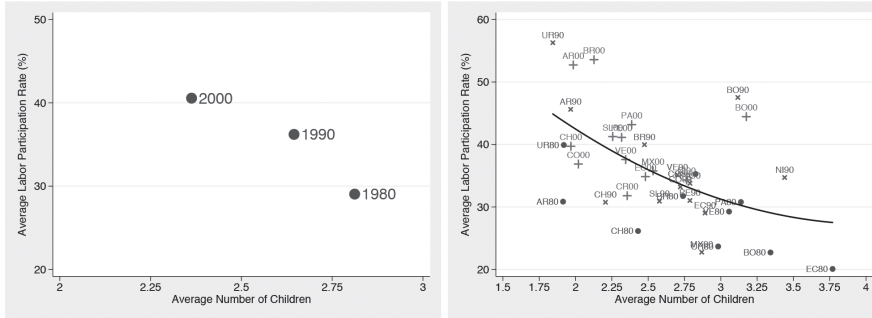
IV. Internal Validity

The analysis in this section deals with the threats to the validity of the identification strategy in its application to Latin American countries. The precondition for the application of the *Same-sex* strategy is the existence of a first-stage relationship between the sex mix of children and further childbearing, which is verified in the next section. However, the crucial point is whether the variation in fertility induced by sex preferences can be considered to be exogenous. Although this cannot be tested formally, the evidence provided here guarantees a greater reliability of the results of the work.

Establishing the case for the randomness of the instrument is relatively straightforward. The gender of a child is a naturally occurring random event, and the sex mix is thus "as good as randomly assigned" (Angrist, 2001). However, a problem arises with extreme forms of son preference that lead to the neglect of daughters in basic healthcare and education, or when sex screening techniques are widely available, and they result in selective abortions and even infanticide (Das Gupta, 2009). In those cases, the sex mix is manipulated and might be correlated with labor supply (the idea that boys contribute relatively more to household welfare compared to girls). The inspection of sex ratios by age and household consumption data can be useful to check if this problem is present in Latin American countries.

Figure Nr. 4 presents the ratio of boys to girls aged zero to four years old for selected countries in 2000. In Latin America and the U.S. the ratios are similar and slightly higher than one, which is an almost universal feature of demographic data. The interesting result is that the ratios in Latin America are substantially lower than in extreme cases like China (CHN), India (IND) and South Korea (KOR). This evidence suggests no discrimination against girls in Latin America in the form of neglected healthcare or feeding those results in higher mortality among girls.

Figure Nr. 4: Sex Ratios -
Number of Boys / Number of Girls, 0 to 4 years old, 2000
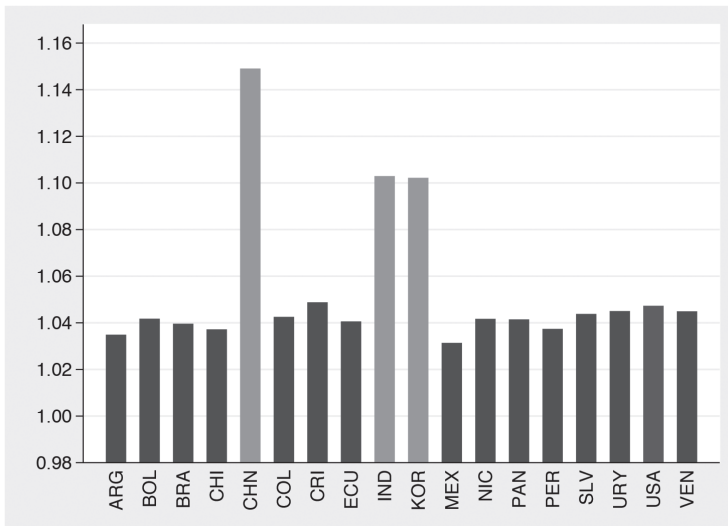
Source: own calculations based on data from United Nations Population Division, World Population Prospects.

Further evidence on the effect of son preferences can be inferred from household's consumption patterns and the budget spent on goods for children of different sex. Table Nr. 1 presents data on budget shares of child-related goods and mean differences by sex composition of children

for Argentina (upper panel) and Colombia (lower panel).[13] If girls were discriminated, parents of boys would spend a higher proportion of their budget on food, health, clothing or education, among other goods. However, in the table none of the differences for parents of two boys or two girls are different from zero at the normal levels of significance.

Another point made in the literature is posed by Rosenzweig and Wolpin (2000) and discussed by CG for Argentina and Mexico. Rosenzweig and Wolpin (2000) argue that in rural India same-sex siblings are related to substantially lower levels of expenditure on child-related goods. These hand-me-down savings could directly affect the marginal utility of leisure and the cost of raising a child, and ultimately the labor supply through mechanisms other than the change in fertility, invalidating the exclusion restriction. Nonetheless, the evidence in Table Nr. 1 shows that the expenditure patterns of households in Argentina and Colombia are not significantly affected by the sex composition of children. Moreover, in the two cases where the difference between budget shares was statistically significant, the sign contradicts the presence of economies of scale, since households whose first two children are girls spend a higher share of their budget on education and clothing.

Finally, while the independence of the sex mix with respect to potential outcomes cannot be established directly, if the instrument is truly random there should not be systematic differences in exogenous characteristics of parents of same-sex and mixed-sex siblings. Using census data this simple check shows that in general women whose first two children were of the same sex and mothers of mixed-sex children cannot be distinguished statistically in terms of age, age at first birth, house ownership and education

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13. Data for Argentina is based on the 2004/2005 Household Expenditure Survey and data for Colombia corresponds to the 2003 Quality of Life Survey, both are nationally representative and were processed following the same criteria as in the census samples. Argentina has two additional categories since it was possible to separate child clothing and education expenditures from adult's expenditures. While it would be desirable to use expenditure surveys for the 14 Latin American countries considered here, household surveys are unusual and also collecting and processing them exceeds the scope of work.

levels.[14] In any case, 2SLS models can accommodate covariates to control for any effect on the outcome of interest.

Table Nr. 1: Differences in budget shares
by sex composition of children

	Share (%)	First two children		
		Same-sex	Two-boys	Two-girls
Argentina - 2004/2005				
Food and beverages	34.0	0.002	-0.000	0.004
		(0.005)	(0.006)	(0.006)
Clothing and footwear	9.4	0.000	-0.004	0.004
		(0.003)	(0.003)	(0.003)
Clothing and footwear (children)	2.9	0.003	0.000	0.004*
		(0.002)	(0.002)	(0.002)
Health and related expenditures	4.9	-0.000	-0.002	0.002
		(0.002)	(0.002)	(0.002)
Education	4.8	0.002	-0.002	0.004***
		(0.001)	(0.002)	(0.002)
Education (children)	2.8	0.005	0.002	0.006
		(0.003)	(0.004)	(0.004)
Colombia - 2003				
Food and beverages	28.8	0.002	0.000	0.003
		(0.004)	(0.005)	(0.005)
Clothing and footwear	7.3	-0.001	-0.003	0.001
		(0.002)	(0.002)	(0.002)
Health and related expenditures	2.1	0.001	0.002	-0.000
		(0.001)	(0.002)	(0.002)
Education	5.0	-0.001	0.000	-0.002
		(0.001)	(0.002)	(0.002)

Note: differences in means (mean of the relevant group minus mean of the rest of the population) and their standard errors (in parentheses). *** significant at 1%, ** significant at 5%, * significant at 10%. The sample consists of 6,815 (Argentina) and 5,825 (Colombia) women aged 18-45 with two or more children aged 18 or younger. Source: Encuesta Nacional de Gasto de los Hogares, INDEC, 2004/2005 (Argentina) and Encuesta de Calidad de Vida, 2003 (Colombia).

14. This was done by running individual regressions of each characteristic on the instrument. The ownership variable is used as a proxy for the household socioeconomic status. The results can be found in the online appendix.

The indirect evidence presented in this section is thus consistent with the internal validity of the Same-sex indicator as an instrument of fertility in the model of labor supply.[15]

V. Results

This section presents the OLS and 2SLS estimates of the relationship between fertility and female labor supply (equation (6)). The analysis was carried out separately on all women and married women. Also, the age of mothers, ages at first birth, sex of the first and second child were included as standard controls.[16] In order to get a visual understanding of the effects over the three decades considered, the coefficients and their confidence intervals are reproduced graphically.[17]

V.1. OLS Estimates

Figure Nr. 5 presents the simple OLS estimates between *Worked for pay* and *More than 2 children* for each country spanning the three decades. In all the cases there is a negative and statistically significant relationship at the 1% level. The effect is relatively constant over time and the magnitude is similar between the countries of Latin America. In Latin America in the year 2000 women with more than two children were on average 11.3 percentage points (p.p.) less likely to participate in the labor force compared to women with only two children, ceteris paribus. In the U.S. that probability is -14.8 p.p. for the year 2000, and it is stronger than in Latin America in all the decades.[18]

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15. *Huber and Mellace (2011) developed a test to assess the validity of an instrumental variable in just-identified models and applied it to the AE's database finding evidence for the validity of Same-sex.*
16. *The over identified model only includes the sex of the first child.*
17. *The tables with all the specifications can be found in the online appendix.*
18. *While useful for comparison purposes, it should be noted that the OLS and IV coefficients are not necessarily comparable since IV estimates identify the causal effect only for the subpopulation of compliers, while OLS coefficient provides a (potentially biased) estimate of the average effect for the whole population.*

Figure Nr. 5: Point estimates and 90% confidence intervals - OLS (all women)
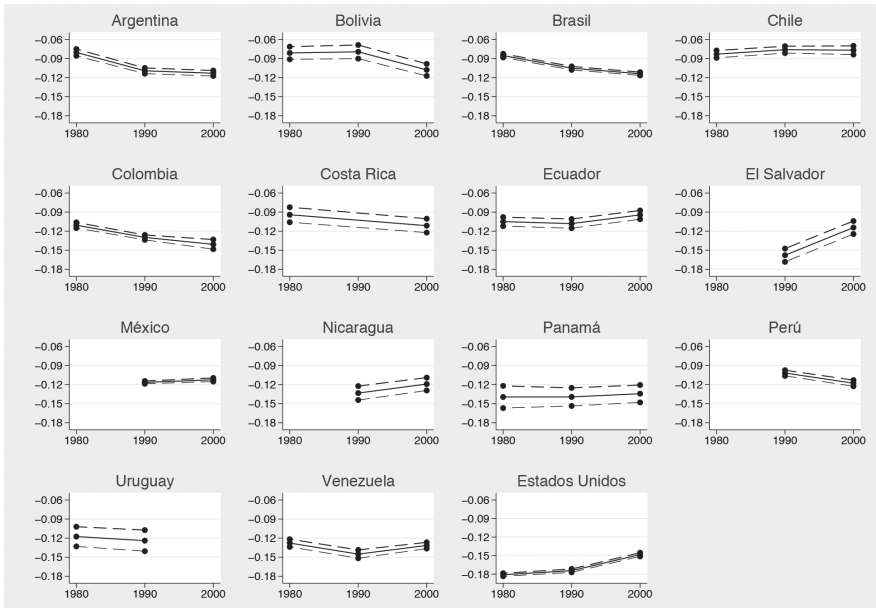
Source: own calculations based on IPUMS-International. Note: the vertical axis shows the coefficient of more than 2 children (estimated by OLS).

V.2. SLS Estimates - First stage

As mentioned in the analysis of the internal validity, a precondition for the application of the *Same-sex* strategy is the existence of a first-stage relationship between the sex mix of children and further childbearing. Figure Nr. 6 summarizes these correlations in all the countries and years considered in the study. Except for Panama 1980, *Same-sex* has a positive and significant effect on fertility at the 1% level. Women with same-sex children are on average more likely to have a third child than women with mixed-sex children, ceteris paribus. However, in Latin America the effect is substantially lower (approximately 3 p.p.) than in the U.S. (approximately 6 p.p.). These results seem to be relatively constant over time.

Regarding the *Two-boys* and *Two-girls* instruments, there is a higher probability of further childbearing for parents of girls (3 p.p.) than for parents of boys (1.9 p.p.). As in AE and CG, this result suggests a moder-

ate bias towards sons which, based on the evidence of the previous section, might be an idiosyncratic feature of the countries under study, and therefore does not constitute a threat to the exogeneity of the instrument. To sum up, these first-stage relationships reveal that a preference for a gender-balanced family (with a moderate bias for sons) is present in Latin American countries, but it is weaker than in the United States.

Figure Nr. 6: First-stage coefficients and 99% confidence intervals - Same-sex (all women)
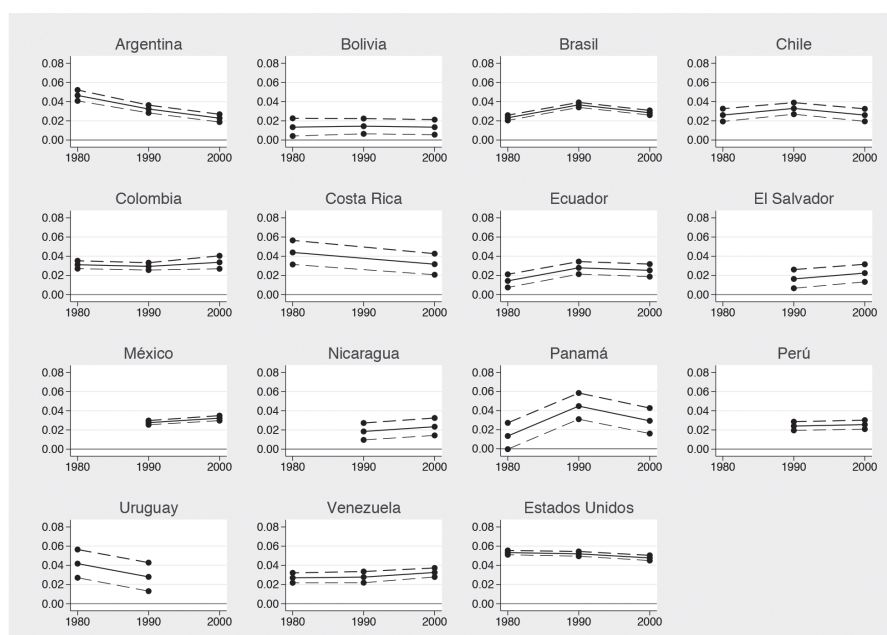
Source: own calculations based on IPUMS-International.
Note: the vertical axis shows the effect of same-sex on more than 2 children.

V.3.SLS Estimates - Second stage

Figure Nr. 7 reports the IV estimates of the effect of fertility on women's labor supply. In general, the point estimates are very imprecise and statistically insignificant. As can be seen from the figure, the 2SLS estimates are centered around the zero-line. The result is robust in nearly all the countries and years and is the main result of the work. There are only three

countries where a significant negative causal effect is found. In the U.S. in 1980 having more than two children reduces women's labor supply in 12 p.p.,[19] 14.5 p.p. in 1990 and 6.1 p.p. in 2000. In the case of Latin American countries the effects are indistinguishable from zero in 1980, while in 1990 Argentina is the only one with a negative and significant effect at 1%. In Mexico 2000 there is a negative effect at the 10% level.[20]

Figure Nr. 7: Point estimates and 90% confidence intervals - IV Same-sex (all women)
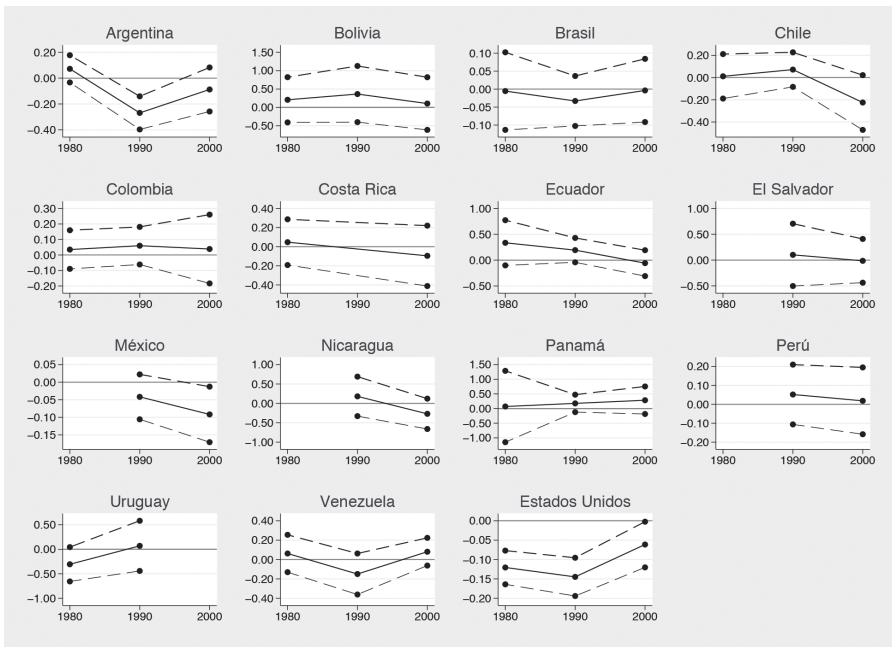
Source: own calculations based on IPUMS-International.
Note: the vertical axis shows the effect of More than 2 children on the probability of working (estimated by 2SLS).

19. The estimated coefficient and standard error in the U.S. coincides exactly with that obtained by AE for 1980.
20. Despite using the same criteria to process the data and using the same variables, the results for Argentina 1991 are twice as high as that of CG, but in Mexico 2000 they are the same. This can be explained as the authors used a sample of 50% for Argentina (599,941 mothers) compared to the 10% (182,824 mothers) used in this work. In the case of Mexico both studies used a sample of 10%.

The results from the over identified model do not differ much from the case in which *Same-sex* is the only instrument. In other words, the use of two instruments improves neither the magnitude nor the accuracy of the results.[21]

The last exercise conducted was to replicate the individual estimates but this time using a pool of mothers for the whole of Latin America in each decade, and also a pool for the three decades altogether. Table Nr. 2 summarizes the results. In the sample of all mothers, OLS estimates agree in sign and magnitude with those found at the country level. Meanwhile, IV estimates are statistically indistinguishable from zero. However, when considering the sample of married mothers, the effect of having more than two children on the labor supply for those women whose fertility decisions are changed by the instrument is negative and statistically significant. The coefficients are approximately half of those estimated by OLS.[22]

Table Nr. 2: OLS and 2SLS estimates of fertility and female labor supply. Pool of countries from Latin America (LA) (all and married (beginning)

	LA 1980	LA 1990	LA 2000	LA 80-90-00	LA 1980	LA 1990	LA 2000	LA 80-90-00
		ALL WOMEN				MARRIED WOMEN		
Panel A – OLS								
More than 2 children	-0.095***	-0.113***	-0.113***	-0.110***	-0.076***	-0.093***	-0.098***	-0.092***
	(0.001)	(0.001)	(0.001)	(0.000)	(0.001)	(0.001)	(0.001)	(0.000)
Panel B - Instrumental Variables								
(1)Same-sex	0.028***	0.029***	0.027***	0.028***	0.030***	0.031***	0.030***	0.030***
	(0.001)	(0.001)	(0.001)	(0.000)	(0.001)	(0.001)	(0.001)	(0.000)
(2)More than 2 children	0.024	-0.029	-0.028	-0.019	-0.027	-0.041*	-0.051**	-0.042***
	(0.033)	(0.022)	(0.025)	(0.015)	(0.030)	(0.021)	(0.023)	(0.014)

21. *In the over identified model the Sargan test allows to analyze whether the results when using Two-boys are statistically different from using Two-girls as the only instrument. This is also a way of checking whether the sex of a child interferes with parent's labor supply for reasons other than family size. In most countries and years considered, the null hypothesis cannot be rejected, evidence for the validity of the instruments.*

22. *Results similar to those obtained for female labor force participation also apply to hours of work (only available in 7 countries) and to an indicator for independent vs dependent work, as dependent variables. The same happens with the number of children as the endogenous independent variable; a sample of mothers with three or more children; heterogeneous effects by education level. The estimates tell essentially the same story and can be checked in the online appendix.*

Table Nr. 2: OLS and 2SLS estimates of fertility and female labor supply.
Pool of countries from Latin America (LA) (all and married (conclusion)

	Panel C - Instrumental Variables – over identified							
(1)Two-boys	0.020***	0.022***	0.019***	0.020***	0.022***	0.024***	0.021***	0.022***
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
(1) Two-girls	0.029***	0.032***	0.030***	0.031***	0.031***	0.034***	0.034***	0.033***
	(0.001)	(0.001)	(0.001)	(0.001)	(0.002)	(0.001)	(0.001)	(0.001)
(2) More than 2 children	0.055	-0.010	-0.013	0.001	-0.005	-0.025	-0.042*	-0.028*
	(0.037)	(0.024)	(0.027)	(0.016)	(0.033)	(0.022)	(0.025)	(0.015)
Sargan p-value	(0.004)	(0.000)	(0.001)	(0.000)	(0.001)	(0.000)	(0.005)	(0.000)
Observations	751,101	1,585,451	1,757,201	4,093,753	686,749	1,436,891	1,508,412	3,632,052
(3) Partial-R2	0.00095	0.00108	0.00087	0.00096	0.00111	0.00126	0.00106	0.00115

Note: Robust standard errors in parentheses. *** significant at 1%, ** significant at 5%, * significant at 10%. Other covariates in the models are Age, Age at first birth, indicators for sex of first and second children, plus year dummies, country dummies and interactions. Samples as described in the text and the appendix. (1) Coefficient of the first stage using More than 2 children as dependent variable; (2) Coefficient of the second stage using Worked for pay as dependent variable; (3) Goodness of t between More than 2 children and Same-sex after controlling for the other covariates in the model.

V.4. Possible explanations

A group of interrelated technical and social factors could be explaining the absence of a causal effect in Latin American countries. The former have to do with weak or low quality instruments, and the latter relate to welfare systems and social norms of Latin American families.

The technical explanation of the imprecise coefficients in the second stage can be attributed to the low quality of the first stage. A usual way of assessing this quality is through two key statistics, the partial-R2 and the F-statistic (Bound, Jaeger and Baker, 1995).[23] In all the countries and years, the F-statistics are all well above the rule of thumb value of 10 which could be due to the use of census data. The remarkable result is that the partial-R2 is extremely

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23. *The partial-R2 isolates the explanatory power of Same-sex over More than 2 children when controlling for age, age at first birth, the sex of the first and second child; the F-statistic in the just-identified model is simply the squared t-statistic, and in the over-identified model, it is a joint significance test of Two-boys and Two-girls. Following Staiger and Stock (1997), it is usual to conclude that an instrument is weak when its F-statistic is lower than 10. However, this is an ad hoc rule and the authors themselves remark that instruments can be weak in large samples even when the statistic is significant at the conventional levels.*

low in all cases, that is, the variability of *Same-sex* explains the variability of fertility very little, and this result is "inherited" by the second stage.[24]

The problem of weak instruments in most countries exacerbates the inherent low precision in the IV estimates compared to OLS, reflected in the large confidence intervals of IV estimates in Figure Nr. 7.[25] In the face of this loss of efficiency, the interpretation of the results calls for caution.

What remains to be clarified is why, despite the low explanatory power of the instrument, there are three countries where a negative causal effect is identified. This result may be due to the magnitude of the *Same-sex* coefficients. Even when the correlations between fertility and *Same-sex* were positive and accurate in all the countries, it is also desirable for the coefficients to be high. For example, if less than 1% of the mothers in the sample had an additional child when the first two were of the same sex, it would be very di cult to detect the effect of that additional child on the labor supply for that group of *compliers*. In the U.S., the coefficients for the first stage (approximately 6 p.p.) were larger than those of Latin America (approximately 3 p.p.). Moreover, even though the F-statistics for Latin American countries were larger than the mentioned threshold, they were notably lower than in the U.S. and the same happened with the partial-R2. This explains why with a stronger first stage in the U.S., it is possible to identify a causal effect. A similar mechanism operates in the case of Argentina and Mexico, although less strongly.

Regarding the estimates with the pool of countries and years, the negative effect in the sample of married women could reflect a less binding budget constraint relative to unmarried women. That is, married women have the option to pool resources with spouses with an income effect that makes their labor supply more elastic, hence adjusting the intensity of their participation.

In the social explanation, a weak first stage for Latin American countries can be related to social and cultural factors. Even though evidence shows

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24. *These results are available in the last two rows of the tables with country level information in the online appendix.*

25. *The variance of the coefficient from the IV estimates can be written as $V(\beta_{IV}) = V(\beta_{OLS}) = R^2_{xz}$, where R^2_{xz} is the partial-R2 corresponding to the first-stage. When the explanatory power of the instrument is low, the standard error of the second stage is magnified (Wooldridge, 2009).*

that the size of Latin American families has fallen in the last decades, the equilibrium size is still above that of developed countries (Chioda, 2011). Then, women with preferences for large families can be indifferent to the sex of their first two children since they may end up having mixed-sex children anyway. In that case, the relevant margin where the mixed-sex sibling preference operates strongly could be at higher parities.[26]

Beyond the problem of weak instruments and the low precision of the estimates, there are other factors that make women labor supply less responsive to fertility in Latin American countries compared to the U.S. First, Latin American households have lower average family income than developed countries. Thus, the economic needs of women to work and supplement those earnings are more binding (Chioda, 2011). Second, in developing countries with large rural sectors it is probably easier for mothers to combine labor and household chores since the physical separation between these activities is smaller than in industrialized countries (Van der Stoep, 2008).[27] Third, since households are larger, older siblings or other relatives are more likely to provide informal childcare freeing up time for mothers to pursue labor market opportunities (Van der Stoep, 2008).[28] Finally, parental leave appears as an important alternative to conciliate family responsibilities with work. While Latin American countries provide on average three paid months of maternity leave (ECLAC, 2011), the U.S. mandates up to 12 weeks of (potentially unpaid) job-protected leave.

VI. Concluding Remarks

This paper studies the causal relationship between fertility and female labor supply in 14 countries in Latin America and the U.S. using census data spanning three decades. The strategy followed builds on an instrumental variable approach, introduced by Angrist and Evans (1998), which relies on parental sex preferences as a source of exogenous variation in fertility.

Even though OLS estimates suggest a negative and statistically signifi-

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26. *In practice this implies working with samples of mothers with three or more children, and thus there is a trade-o between stronger first-stages and smaller samples.*
27. *This may not be the case here because the indicator Worked for pay rules out unpaid family work.*
28. *In developed countries this could be compensated by the broader access to formal childcare like nurseries and daycare centers.*

cant relationship between fertility and mothers' labor supply in each Latin American country, 2SLS estimates fail to identify a causal effect in most of them. Namely, the average effect of moving from a two-child family to a larger one is statistically zero for those women whose fertility decisions are changed by the instrument (*compliers*). The results for the U.S., Argentina and Mexico agree with those reported in previous studies. When considering a pool of countries in Latin America, a negative and statistically significant causal effect is found in the sample of married women.

Despite having highly accurate first-stage estimates (and indirect evidence consistent with the internal validity of the instrument), the group of compliers is small compared to that of developed countries like the U.S. Moreover, the analysis of the quality of the instrument reveals a weak explanatory power of sibling sex composition on fertility. The problem of weak instruments entails a loss of precision in the second-stage, with large standard errors that make the interpretation of IV estimates meaningless.

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Appendix

Table A.1: Countries, years and official source of census

Country		Year of Census	Fraction (%)	Census agency
Argentina	ARG	1980, 1991, 2001	10	Argentine National Institute of Statistics and Censuses (INDEC)
Bolivia	BOL	1976, 1992, 2001	10	National Institute of Statistics, Ministry of Planning and Coordination
Brazil	BRA	1980, 1991, 2000	5-6	Instituto Brasileiro de Geografia e Estatistica
Chile	CHL	1982, 1992, 2002	10	Instituto Nacional de Estadísticas
Colombia	COL	1985, 1993, 2005	10	Departamento Administrativo Nacional de Estadística (DANE)
Costa Rica	CRI	1984, 2000	10	Instituto Nacional de Estadística y Censos
Ecuador	ECU	1982, 1990, 2001	10	Instituto Nacional de Estadística y Censos
El Salvador	SLV	1992, 2007	10	General Directorate of Statistics and Censuses
Mexico	MEX	1990, 2000	10-10.6	Instituto Nacional de Estadística, Geografía e Informática (INEGI)
Nicaragua	NIC	1995, 2005	10	National Institute of Statistics and Censuses
Panama	PAN	1980, 1990, 2000	10	Contraloria General de la Republica, Direccion de Estadistica y Censo
Peru	PER	1993, 2007	10	National Institute of Statistics and Computing
Uruguay	URY	1985, 1996	10	National Institute of Statistics
Venezuela	VEN	1981, 1990, 2001	10	O cina Central de Estadística e Informática
United States	USA	1980, 1990, 2000	5	U.S. Census Bureau

Source: IPUMS-International.

A.2. Matching between mothers and children

The study of female labor supply and fertility required a subsample of mothers with the characteristics of their children linked to them. Other studies had to restrict the sample to women who are heads of household or spouses of the head since they only had access to a variable that describes the relationship of the individual to the head of household. In this work, I use the variable momloc, which indicates whether or not the person's mother lived in the same household and, if so, it gives the person number of the mother. Thus, it was easier to link the characteristics of children and their mothers, and for those households with more than one mother it was possible to identify each son separately obtaining larger samples than the other studies.

Since the identification strategy uses the Same-sex indicator for whether the sex of the second child matches the sex of the first child, the sample

was restricted to mothers with at least two children. In addition, and as in other studies with similar settings, the sample was restricted to women aged 21 to 35 years, whose first child was under 18 years old, and whose second child was at least 1 year old. On the one hand, this is because few women under 21 have two children, and on the other hand, it is common for children over 18 years to leave their parents' house for educational or employment purposes, and censuses do not keep track of the children as they become independent. Restriction to mothers aged 35 years or less means that the cut to age at 18 for the oldest child does not generate a highly selected sample. Finally, the observations on which the age of the mother and spouse at the birth of the first child was under 14 were discarded, since these cases may constitute data entry errors.

A.3. A note on weighting

Solon, Haider and Wooldridge (2013) discuss the use of sampling weights in two situations: to estimate population descriptive statistics and to estimate causal effects. The first situation is obvious when someone is working with population samples. In the second case, the motive for weighting is to achieve consistent estimates by correcting for endogenous sampling. If the sample is systematically unrepresentative of the population in a known manner (e.g. oversampling of low-income households in the context of Mincer regressions) the estimates could be inconsistent. The sampling would be endogenous because the sampling criterion, i.e. the population income, is correlated with the error term in the regression. When performing IV estimates, one would need to weight the orthogonality conditions by the inverse probabilities of selection in order to get consistent estimates. However, an important point stressed in Wooldridge (1999) is that, if the sampling probabilities vary exogenously instead of endogenously, weighting might be unnecessary for consistency and harmful for precision.[29]

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29. *Solon et al. (2013) provide an example in which a linear regression model, including state dummy variables among the explanatory variables, is estimated, with a U.S. sample that over-represents certain states (as in the Current Population Survey). Then, if the model is correctly specified, the error term is not related to the sampling criterion and weighting is unnecessary. Moreover, if the error term were homoscedastic prior to weighting, the weighting would induce heteroscedasticity and imprecise estimates.*

In the present work, most of IPUMS samples are "at", i.e., each person in the sample represents a fixed number of people in the population. Only 30% of the samples used have weights with some records representing more cases than others. This means that some people with certain characteristics are overrepresented in the samples, and others are underrepresented. In all cases, the databases provide the harmonized variable WTPER. For instance, in Argentina 2001 all the records have a weighting of 10%, and in 1980 and 1990 the weight varies within the sample. Then, it is important to include sampling weights in the regressions at the country and year level. However, when using the pool of mothers for the whole of Latin America, the inclusion of sampling weights is less obvious since they are defined within the country but not between countries. In this case, country and year dummies were included as control variables.

Table A.2: Descriptive Statistics – 1980s

	Argentina	Bolivia	Brazil	Chile	Colombia	C. Rica	Ecuador	Panama	Uruguay	Venez.	LA	U.S.
Worked for pay	0.20	0.15	0.20	0.16	0.26	0.18	0.14	0.27	0.33	0.23	0.20	0.46
(=1 if worked for pay)	(0.40)	(0.36)	(0.40)	(0.36)	(0.44)	(0.38)	(0.34)	(0.44)	(0.47)	(0.42)	(0.40)	(0.50)
More than 2 children	0.53	0.70	0.66	0.51	0.59	0.59	0.69	0.68	0.47	0.67	0.63	0.41
(=1 if had more than 2)	(0.50)	(0.46)	(0.47)	(0.50)	(0.49)	(0.49)	(0.46)	(0.46)	(0.50)	(0.47)	(0.48)	(0.49)
Number of children	2.9	3.3	3.4	2.8	3.1	3.1	3.4	3.4	2.8	3.4	3.3	2.6
	(1.2)	(1.3)	(1.5)	(1.1)	(1.3)	(1.3)	(1.4)	(1.4)	(1.1)	(1.5)	(1.5)	(0.9)
Age	29.6	29.2	29.2	29.6	29.1	29.0	28.9	29.1	29.9	28.8	29.2	30.1
	(3.8)	(3.8)	(3.9)	(3.8)	(3.9)	(3.9)	(3.9)	(3.9)	(3.7)	(3.9)	(3.9)	(3.6)
Age at 1st birth	21.5	20.5	20.7	20.5	20.1	20.1	20.1	19.8	21.1	20.0	20.6	20.8
	(3.5)	(3.2)	(3.2)	(3.2)	(3.2)	(3.1)	(3.2)	(3.1)	(3.3)	(3.3)	(3.3)	(3.1)
Same-sex	0.508	0.498	0.497	0.509	0.510	0.502	0.502	0.506	0.508	0.507	0.502	0.504
(=1 if same-sex siblings)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)
Two-boys	0.258	0.255	0.254	0.261	0.261	0.260	0.261	0.267	0.262	0.263	0.257	0.261
(=1 if both were boys)	(0.437)	(0.436)	(0.436)	(0.439)	(0.439)	(0.438)	(0.439)	(0.442)	(0.440)	(0.440)	(0.437)	(0.439)
Two-girls	0.247	0.241	0.238	0.245	0.243	0.238	0.239	0.236	0.243	0.238	0.240	0.240
(=1 if both were girls)	(0.431)	(0.428)	(0.426)	(0.430)	(0.429)	(0.426)	(0.427)	(0.424)	(0.429)	(0.426)	(0.427)	(0.427)
Boy 1st	0.509	0.504	0.511	0.514	0.512	0.511	0.511	0.516	0.509	0.513	0.511	0.512
(=1 if first was a boy)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)
Boy 2nd	0.502	0.509	0.505	0.504	0.507	0.511	0.511	0.516	0.510	0.512	0.506	0.510
(=1 if second was a boy)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)
Education (years)	7.13	2.64	3.54	7.48	5.37	6.62	5.33	6.43	8.12	5.79	4.72	11.96
	(3.92)	(3.35)	(3.47)	(3.73)	(3.62)	(3.60)	(4.01)	(3.97)	(3.88)	(3.54)	(3.89)	(2.41)
Observations	122,344	22,524	279,047	50,984	130,980	14,147	40,927	9,996	10,782	72,966	754,697	476,582

Source: own calculations based on IPUMS-International. Note: mean and standard deviation (in parentheses). Samples correspond to women aged 21-35 years with two or more children.

Tabla A.3: Descriptive Statistics - 1990s

	Argentina	Bolivia	Brasil	Chile	Colombia	Ecuador	El Salv.	Mexico	Nicarag.	Panam	Peru	Uruguay	Venez.	LA	U.S.
Worked for pay	0.33	0.32	0.30	0.19	0.25	0.23	0.29	0.15	0.28	0.25	0.17	0.44	0.27	0.25	0.55
(=1 if worked for pay)	(0.47)	(0.47)	(0.46)	(0.39)	(0.43)	(0.42)	(0.45)	(0.35)	(0.45)	(0.43)	(0.38)	(0.50)	(0.44)	(0.43)	(0.50)
More than 2 children	0.57	0.68	0.54	0.42	0.53	0.62	0.61	0.66	0.69	0.60	0.61	0.44	0.61	0.58	0.40
(=1 if had more than 2)	(0.49)	(0.47)	(0.50)	(0.49)	(0.50)	(0.48)	(0.49)	(0.48)	(0.46)	(0.49)	(0.49)	(0.50)	(0.49)	(0.49)	(0.49)
Number of children	3.0	3.3	2.9	2.6	2.9	3.2	3.1	3.3	3.5	3.1	3.1	2.7	3.2	3.1	2.6
	(1.2)	(1.3)	(1.2)	(0.8)	(1.1)	(1.3)	(1.2)	(1.4)	(1.5)	(1.2)	(1.2)	(1.1)	(1.3)	(1.3)	(0.8)
Age	29.8	29.4	29.6	30.0	29.5	29.3	29.1	29.3	28.5	29.2	29.4	30.2	29.4	29.5	30.4
	(3.8)	(3.8)	(3.8)	(3.6)	(3.8)	(3.8)	(3.9)	(3.8)	(4.1)	(3.9)	(3.8)	(3.7)	(3.9)	(3.8)	(3.5)
Age at 1st birth	20.8	20.3	20.5	20.8	20.1	20.1	19.6	19.8	19.1	19.9	20.3	21.0	20.0	20.3	21.4
	(3.4)	(3.1)	(3.2)	(3.3)	(3.2)	(3.2)	(3.1)	(3.1)	(3.0)	(3.2)	(3.2)	(3.4)	(3.2)	(3.2)	(3.6)
Same-sex	0.508	0.497	0.492	0.504	0.510	0.504	0.509	0.503	0.508	0.509	0.505	0.502	0.506	0.500	0.504
(=1 if same-sex siblings)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)
Two-boys	0.257	0.254	0.255	0.262	0.263	0.263	0.282	0.257	0.262	0.266	0.256	0.254	0.260	0.258	0.261
(=1 if both were boys)	(0.437)	(0.436)	(0.436)	(0.440)	(0.440)	(0.440)	(0.450)	(0.437)	(0.440)	(0.442)	(0.436)	(0.435)	(0.438)	(0.437)	(0.439)
Two-girls	0.248	0.240	0.234	0.238	0.244	0.239	0.224	0.243	0.241	0.239	0.247	0.246	0.240	0.239	0.239
(=1 if both were girls)	(0.432)	(0.427)	(0.423)	(0.426)	(0.429)	(0.426)	(0.417)	(0.429)	(0.428)	(0.427)	(0.431)	(0.430)	(0.427)	(0.427)	(0.427)
Boy 1st	0.505	0.507	0.512	0.511	0.513	0.513	0.543	0.509	0.514	0.515	0.505	0.508	0.515	0.511	0.513
(=1 if first was a boy)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.498)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)
Boy 2nd	0.505	0.507	0.509	0.513	0.506	0.511	0.514	0.506	0.507	0.512	0.503	0.499	0.505	0.507	0.510
(=1 if second was a boy)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)
Education (years)	8.5	5.5	5.2	8.9	6.0	6.9	4.8	5.7	4.9	8.0	6.5	8.1	6.7	6.0	.
	(3.9)	(4.4)	(3.9)	(3.4)	(3.4)	(4.5)	(4.3)	(3.8)	(4.1)	(4.1)	(4.1)	(3.1)	(3.8)	(4.0)	.
Observations	183,366	27,899	432,303	63,097	154,817	48,737	24,090	416,856	25,035	11,691	102,864	10,410	89,508	1,590,673	481,039

Source: own calculations based on IPUMS-International. Note: mean and standard deviation (in parentheses). Samples correspond to women aged 21-35 years with two or more children.

Tabla A.4: Descriptive Statistics - 2000s

	Argent.	Bolivia	Brasil	Chile	Colomb.	C. Rica	Ecuador	El Salv.	Mexico	Nicarag.	Panam	Peru	Venez.	LA	U.S.
Worked for pay	0.28	0.38	0.38	0.27	0.31	0.24	0.28	0.36	0.27	0.32	0.27	0.28	0.32	0.32	0.58
(=1 if worked for pay)	(0.45)	(0.49)	(0.48)	(0.44)	(0.46)	(0.43)	(0.45)	(0.48)	(0.44)	(0.47)	(0.45)	(0.45)	(0.47)	(0.47)	(0.49)
More than 2 children	0.55	0.62	0.46	0.34	0.47	0.52	0.53	0.50	0.55	0.56	0.55	0.46	0.54	0.50	0.42
(=1 if had more than 2)	(0.50)	(0.49)	(0.50)	(0.48)	(0.50)	(0.50)	(0.50)	(0.50)	(0.50)	(0.50)	(0.50)	(0.50)	(0.50)	(0.50)	(0.49)
Number of children	3.0	3.1	2.7	2.5	2.7	2.8	2.9	2.8	2.9	3.0	3.0	2.7	2.9	2.8	2.6
	(1.3)	(1.2)	(1.0)	(0.7)	(1.0)	(1.0)	(1.1)	(1.0)	(1.1)	(1.2)	(1.2)	(1.0)	(1.1)	(1.1)	(0.9)
Age	29.7	29.3	29.6	30.6	29.6	29.9	29.3	29.7	29.5	28.9	29.5	29.9	29.5	29.6	30.4
	(3.8)	(3.9)	(3.9)	(3.4)	(3.9)	(3.8)	(3.9)	(3.7)	(3.8)	(4.0)	(3.9)	(3.7)	(3.9)	(3.8)	(3.6)
Age at 1st birth	20.3	20.2	20.1	20.6	19.9	20.0	19.9	19.5	20.1	19.0	20.0	20.1	19.8	20.1	21.5
	(3.4)	(3.1)	(3.1)	(3.2)	(3.1)	(3.1)	(3.2)	(3.1)	(3.1)	(2.9)	(3.3)	(3.2)	(3.2)	(3.2)	(3.7)
Same-sex	0.521	0.504	0.501	0.502	0.502	0.503	0.508	0.509	0.502	0.508	0.500	0.507	0.504	0.504	0.505
(=1 if same-sex siblings)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)
Two-boys	0.267	0.262	0.259	0.260	0.262	0.262	0.262	0.258	0.258	0.261	0.261	0.262	0.260	0.260	0.261
(=1 if both were boys)	(0.443)	(0.440)	(0.438)	(0.439)	(0.440)	(0.439)	(0.440)	(0.438)	(0.438)	(0.439)	(0.439)	(0.440)	(0.439)	(0.439)	(0.439)
Two-girls	0.251	0.239	0.239	0.239	0.237	0.238	0.243	0.246	0.241	0.242	0.236	0.242	0.240	0.241	0.240
(=1 if both were girls)	(0.433)	(0.427)	(0.426)	(0.427)	(0.425)	(0.426)	(0.429)	(0.431)	(0.428)	(0.429)	(0.424)	(0.428)	(0.427)	(0.427)	(0.427)
Boy 1st	0.508	0.510	0.512	0.514	0.515	0.512	0.512	0.508	0.510	0.515	0.517	0.510	0.512	0.511	0.512
(=1 if rst was a boy)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)
Boy 2nd	0.508	0.512	0.508	0.507	0.510	0.511	0.507	0.505	0.507	0.505	0.508	0.510	0.508	0.508	0.509
(=1 if second was a boy)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)
Education (years)	9.0	6.6	5.9	10.1	7.3	7.3	7.6	6.2	7.4	5.4	8.3	7.8	7.8	7.0	.
	(3.6)	(4.4)	(3.7)	(3.5)	(4.0)	(3.5)	(4.4)	(4.5)	(3.8)	(4.1)	(4.0)	(3.9)	(3.2)	(3.9)	.
Observations	138,321	34,436	459,385	52,009	167,541	19,137	51,887	27,302	559,994	26,944	13,016	106,982	105,690	1,762,644	423,549

Source: own calculations based on IPUMS-International. Note: mean and standard deviation (in parentheses). Samples correspond to women aged 21–35 years with two or more children.

Table A.5: OLS and 2SLS estimates of fertility and female labor supply. 1980s census (ALL WOMEN)

	ARG	BOL	BRA	CHL	COL	CRI	ECU	PAN	URY	VEN	USA
Panel A - OLS											
More than 2 children	-0.080***	-0.081***	-0.085***	-0.083***	-0.111***	-0.094***	-0.105***	-0.139***	-0.118***	-0.128***	-0.181***
	(0.003)	(0.006)	(0.002)	(0.004)	(0.003)	(0.007)	(0.004)	(0.011)	(0.009)	(0.004)	(0.001)
Panel B - Instrumental Variables											
(1) Same-sex	0.046***	0.013***	0.023***	0.026***	0.031***	0.044***	0.014***	0.013	0.042***	0.027***	0.053***
	(0.003)	(0.006)	(0.002)	(0.004)	(0.003)	(0.008)	(0.004)	(0.008)	(0.009)	(0.003)	(0.001)
(2) More than 2 children	0.072	0.208	-0.005	0.012	0.035	0.047	0.336	0.070	-0.306	0.062	-0.120***
	(0.064)	(0.374)	(0.066)	(0.122)	(0.076)	(0.145)	(0.266)	(0.740)	(0.213)	(0.117)	(0.026)
Panel C - Instrumental Variables – over identified											
(1) Two-boys	0.029***	0.016**	0.018***	0.012**	0.027***	0.037***	0.010*	0.006	0.027**	0.020***	0.041***
	(0.005)	(0.008)	(0.002)	(0.006)	(0.004)	(0.011)	(0.006)	(0.012)	(0.013)	(0.004)	(0.002)
(1) Two-girls	0.057***	0.010	0.022***	0.033***	0.028***	0.045***	0.016***	0.017	0.051***	0.027***	0.058***
	(0.005)	(0.008)	(0.002)	(0.006)	(0.004)	(0.011)	(0.006)	(0.012)	(0.013)	(0.005)	(0.002)
(2) More than 2 children	0.088	0.241	0.032	-0.037	0.061	0.046	0.392	0.344	-0.289	0.100	-0.118***
	(0.066)	(0.389)	(0.076)	(0.131)	(0.086)	(0.157)	(0.302)	(0.811)	(0.219)	(0.136)	(0.028)
Sargan p-value	(0.746)	(0.792)	(0.044)	(0.288)	(0.026)	(0.256)	(0.356)	(0.515)	(0.694)	(0.262)	(0.058)
Observations	121,832	22,492	277,606	50,779	130,265	14,092	40,828	9,965	10,751	72,491	474,846
(3) Parcial-R2	0.0025	0.0003	0.0007	0.0008	0.0012	0.0023	0.0003	0.0002	0.0021	0.0009	0.0032
(4) F-statistic	180.7	5.7	193.6	41.7	159.2	33.1	11.8	2.0	22.3	68.8	1,548.3

Note: Robust standard errors in parentheses. *** significant at 1%, ** significant at 5%, * significant at 10%. Other covariates in the models are Age, Age at first birth, plus indicators for the sex of the first and second children. Samples as described in the text and the appendix. All models include sample weights (see appendix A.3). (1) Coefficient of the first stage using More than 2 children as dependent variable; (2) Coefficient of the second stage using Worked for pay as dependent variable; (3) Goodness of t between More than 2 children and Same-sex after controlling for the other covariates in the model; (4) F-statistic on Same-sex in the first stage regression

Table A.6. OLS and 2SLS estimates of fertility and female labor supply. 1980s census (MARRIED WOMEN)

	ARG	BOL	BRA	CHL	COL	CRI	ECU	PAN	URY	VEN	USA
Panel A - OLS											
More than 2 children	-0.065***	-0.056***	-0.066***	-0.065***	-0.089***	-0.076***	-0.092***	-0.125***	-0.114***	-0.100***	-0.168***
	(0.003)	(0.006)	(0.002)	(0.003)	(0.003)	(0.007)	(0.004)	(0.011)	(0.010)	(0.004)	(0.002)
Panel B - Instrumental Variables											
(1) Same-sex	0.050***	0.012**	0.025***	0.027***	0.033***	0.047***	0.018***	0.016*	0.048***	0.029***	0.058***
	(0.004)	(0.006)	(0.002)	(0.004)	(0.003)	(0.008)	(0.004)	(0.009)	(0.009)	(0.003)	(0.001)
(2) More than 2 children	0.030	0.079	-0.046	0.045	-0.060	0.048	0.321	-0.318	-0.315*	-0.063	-0.120***
	(0.058)	(0.407)	(0.060)	(0.115)	(0.072)	(0.130)	(0.208)	(0.628)	(0.191)	(0.103)	(0.027)
Panel C - Instrumental Variables – over identified											
(1) Two-boys	0.033***	0.014*	0.019***	0.014**	0.028***	0.041***	0.012**	0.015	0.032***	0.025***	0.045***
	(0.005)	(0.008)	(0.002)	(0.006)	(0.004)	(0.011)	(0.006)	(0.012)	(0.013)	(0.005)	(0.002)
(1) Two-girls	0.061***	0.009	0.024***	0.034***	0.031***	0.047***	0.022***	0.014	0.057***	0.028***	0.064***
	(0.005)	(0.008)	(0.002)	(0.006)	(0.004)	(0.012)	(0.006)	(0.013)	(0.013)	(0.005)	(0.002)
(2) More than 2 children	0.039	0.217	-0.019	0.013	-0.050	0.052	0.364	-0.345	-0.262	-0.043	-0.118***
	(0.060)	(0.435)	(0.068)	(0.124)	(0.080)	(0.138)	(0.224)	(0.704)	(0.195)	(0.116)	(0.028)
Sargan p-value	(0.828)	(0.187)	(0.011)	(0.306)	(0.171)	(0.376)	(0.481)	(0.214)	(0.195)	(0.060)	(0.090)
Observations	114,575	20,794	262,814	46,504	116,481	12,305	37,737	8,604	9,913	63,176	398,171
(3) Partial-R2	0.0029	0.0002	0.0008	0.0009	0.0014	0.0027	0.0005	0.0003	0.0027	0.0012	0.0038
(4) F-statistic	197.2	4.2	207.5	41.1	159.7	33.1	17.8	2.9	26.9	75.9	1,529.3

Note: Robust standard errors in parentheses. *** significant at 1%, ** significant at 5%, * significant at 10%. Other covariates in the models are Age, Age at first birth, plus indicators for the sex of the first and second children. Samples as described in the text and the appendix. All models include sample weights (see appendix A.3). (1) Coefficient of the first stage using More than 2 children as dependent variable; (2) Coefficient of the second stage using Worked for pay as dependent variable; (3) Goodness of t between More than 2 children and Same-sex after controlling for the other covariates in the model; (4) F-statistic on Same-sex in the first stage regression.

Table A.7: OLS and 2SLS estimates of fertility and female labor supply. 1990s census (ALL WOMEN)

	ARG	BOL	BRA	CHL	COL	ECU	SLV	MEX	NIC	PAN	PER	URY	VEN	USA
Panel A - OLS														
More than 2 children	-0.109***	-0.079***	-0.105***	-0.076***	-0.130***	-0.108***	-0.158***	-0.116***	-0.133***	-0.139***	-0.102***	-0.124***	-0.145***	-0.174***
	(0.003)	(0.007)	(0.002)	(0.003)	(0.002)	(0.004)	(0.006)	(0.001)	(0.007)	(0.009)	(0.003)	(0.010)	(0.004)	(0.002)
Panel B - Instrumental Variables														
(1) Same-sex	0.032***	0.014***	0.037***	0.033***	0.029***	0.028***	0.016***	0.028***	0.018***	0.045***	0.024***	0.028***	0.028***	0.052***
	(0.002)	(0.005)	(0.002)	(0.004)	(0.002)	(0.004)	(0.006)	(0.001)	(0.005)	(0.008)	(0.003)	(0.009)	(0.004)	(0.002)
(2) More than 2 children	-0.269***	0.364	-0.033	0.072	0.060	0.194	0.102	-0.042	0.181	0.176	0.052	0.070	-0.149	-0.145***
	(0.078)	(0.465)	(0.042)	(0.095)	(0.074)	(0.143)	(0.367)	(0.039)	(0.310)	(0.181)	(0.096)	(0.311)	(0.129)	(0.030)
Panel C - Instrumental Variables – over identified														
(1) Two-boys	0.019***	0.010	0.027***	0.032***	0.024***	0.024***	0.013*	0.019***	0.020***	0.047***	0.024***	0.008	0.023***	0.048***
	(0.003)	(0.007)	(0.002)	(0.005)	(0.003)	(0.006)	(0.008)	(0.002)	(0.007)	(0.012)	(0.004)	(0.013)	(0.005)	(0.002)
(1) Two-girls	0.041***	0.017***	0.041***	0.026***	0.028***	0.029***	0.015*	0.033***	0.011	0.037***	0.021***	0.042***	0.024***	0.049***
	(0.003)	(0.007)	(0.002)	(0.005)	(0.003)	(0.006)	(0.009)	(0.002)	(0.008)	(0.012)	(0.004)	(0.013)	(0.005)	(0.002)
(2) More than 2 children	-0.272***	0.433	0.001	0.057	0.081	0.242	0.104	-0.026	0.136	0.169	0.055	0.247	-0.153	-0.151***
	(0.079)	(0.517)	(0.045)	(0.105)	(0.084)	(0.154)	(0.425)	(0.040)	(0.340)	(0.191)	(0.103)	(0.306)	(0.155)	(0.032)
Sargan p-value	(0.438)	(0.710)	(0.000)	(0.074)	(0.030)	(0.116)	(0.168)	(0.033)	(0.473)	(0.592)	(0.827)	(0.372)	(0.104)	(0.986)
Observations	183,874	27,950	433,696	63,307	155,386	48,847	24,170	417,896	25,165	11,737	103,092	10,441	90,101	482,548
(3) Parcial-R2	0.0012	0.0002	0.0016	0.0013	0.0010	0.0010	0.0003	0.0011	0.0005	0.0024	0.0008	0.0011	0.0009	0.0031
(4) F-statistic	172.7	6.2	549.7	80.1	162.1	47.3	7.4	438.6	12.2	28.5	78.8	11.4	58.8	1,178.6

*Note: Robust standard errors in parentheses. *** significant at 1%, ** significant at 5%, * significant at 10%. Other covariates in the models are Age, Age at first birth, plus indicators for the sex of the first birth and second children. Samples as described in the text and the appendix. All models include sample weights (see appendix A.3). (1) Coefficient of the first stage using More than 2 children as dependent variable; (2) Coefficient of the second stage using Worked for pay as dependent variable; (3) Goodness of t between More than 2 children and Same-sex after controlling for the other covariates in the model; (4) F-statistic on Same-sex in the first stage regression.*

Table A.8: OLS and 2SLS estimates of fertility and female labor supply. 1990s census (MARRIED WOMEN)

	ARG	BOL	BRA	CHL	COL	ECU	SLV	MEX	NIC	PAN	PER	URY	VEN	USA
Panel A - OLS														
More than 2 children	-0.092***	-0.058***	-0.086***	-0.058***	-0.109***	-0.093***	-0.140***	-0.091***	-0.121***	-0.132***	-0.088***	-0.126***	-0.126***	-0.166***
	(0.003)	(0.007)	(0.002)	(0.003)	(0.002)	(0.005)	(0.007)	(0.001)	(0.007)	(0.009)	(0.003)	(0.011)	(0.004)	(0.002)
Panel B - Instrumental Variables														
(1) Same-sex	0.033***	0.015***	0.039***	0.039***	0.033***	0.028***	0.026***	0.029***	0.022***	0.048***	0.027***	0.029***	0.030***	0.060***
	(0.003)	(0.005)	(0.002)	(0.004)	(0.002)	(0.004)	(0.006)	(0.001)	(0.006)	(0.009)	(0.003)	(0.010)	(0.004)	(0.002)
(2) More than 2 children	-0.208***	0.185	-0.059	-0.001	0.074	0.328**	-0.038	-0.058*	0.076	0.110	-0.020	0.154	-0.003	-0.157***
	(0.077)	(0.428)	(0.040)	(0.078)	(0.066)	(0.154)	(0.218)	(0.035)	(0.263)	(0.173)	(0.086)	(0.306)	(0.121)	(0.029)
Panel C - Instrumental Variables – over identified														
(1) Two-boys	0.019***	0.009	0.029***	0.037***	0.028***	0.023***	0.015*	0.020***	0.023***	0.047***	0.026***	0.009	0.027***	0.055***
	(0.004)	(0.007)	(0.002)	(0.005)	(0.003)	(0.006)	(0.009)	(0.002)	(0.008)	(0.013)	(0.004)	(0.013)	(0.005)	(0.002)
(1) Two-girls	0.042***	0.019***	0.044***	0.033***	0.032***	0.030***	0.034***	0.034***	0.015*	0.043***	0.024***	0.044***	0.025***	0.058***
	(0.004)	(0.007)	(0.002)	(0.006)	(0.004)	(0.006)	(0.009)	(0.002)	(0.008)	(0.013)	(0.004)	(0.014)	(0.006)	(0.002)
(2) More than 2 children	-0.215***	0.234	-0.029	-0.012	0.088	0.402***	-0.042	-0.041	0.048	0.108	-0.021	0.245	0.038	-0.164***
	(0.078)	(0.454)	(0.042)	(0.086)	(0.073)	(0.167)	(0.228)	(0.037)	(0.286)	(0.184)	(0.092)	(0.302)	(0.142)	(0.031)
Sargan p-value	(0.644)	(0.881)	(0.000)	(0.083)	(0.064)	(0.091)	(0.918)	(0.009)	(0.541)	(0.641)	(0.717)	(0.676)	(0.075)	(0.487)
Observations	168,990	26,035	387,639	56,771	134,890	44,720	20,513	394,750	20,510	9,954	94,911	9,260	75,631	380,951
(3) Parcial-R2	0.0013	0.0003	0.0018	0.0017	0.0013	0.0010	0.0009	0.0012	0.0007	0.0028	0.0009	0.0013	0.0011	0.0041
(4) F-statistic	167.2	6.6	565.8	98.6	181.3	42.8	17.6	455.3	14.9	28.0	90.0	12.0	61.5	1,266.6

Note: Robust standard errors in parentheses. *** significant at 1%, ** significant at 5%, * significant at 10%. Other covariates in the models are Age, Age at first birth, plus indicators for the sex of the first and second children. Samples as described in the text and the appendix. All models include sample weights (see appendix A.3). (1) Coefficient of the first stage using More than 2 children as dependent variable; (2) Coefficient of the second stage using Worked for pay as dependent variable; (3) Goodness of t between More than 2 children and Same-sex after controlling for the other covariates in the model; (4) F-statistic on Same-sex in the first stage regression.

Table A.9. OLS and 2SLS estimates of fertility and female labor supply. 2000s census (ALL WOMEN)

	ARG	BOL	BRA	CHL	COL	CRI	ECU	SLV	MEX	NIC	PAN	PER	VEN	USA
Panel A - OLS														
More than 2 children	-0.113***	-0.108***	-0.114***	-0.077***	-0.141***	-0.111***	-0.094***	-0.114***	-0.112***	-0.119***	-0.134***	-0.118***	-0.132***	-0.148***
	(0.003)	(0.006)	(0.002)	(0.004)	(0.005)	(0.007)	(0.004)	(0.006)	(0.002)	(0.006)	(0.008)	(0.003)	(0.003)	(0.002)
Panel B - Instrumental Variables														
(1) Same-sex	0.023***	0.013***	0.028***	0.026***	0.034***	0.032***	0.025***	0.022***	0.032***	0.023***	0.030***	0.026***	0.033***	0.048***
	(0.002)	(0.005)	(0.001)	(0.004)	(0.004)	(0.007)	(0.004)	(0.006)	(0.002)	(0.006)	(0.008)	(0.003)	(0.003)	(0.002)
(2) More than 2 children	-0.088	0.105	-0.004	-0.224	0.039	-0.096	-0.058	-0.014	-0.092*	-0.269	0.300	0.018	0.081	-0.061*
	(0.103)	(0.435)	(0.053)	(0.150)	(0.135)	(0.192)	(0.152)	(0.258)	(0.048)	(0.238)	(0.285)	(0.107)	(0.087)	(0.036)
Panel C - Instrumental Variables – over identified														
(1) Two-boys	0.013***	0.007	0.022***	0.018***	0.028***	0.030***	0.019***	0.007	0.021***	0.014*	0.014	0.016***	0.030***	0.040***
	(0.003)	(0.007)	(0.002)	(0.006)	(0.006)	(0.009)	(0.006)	(0.008)	(0.002)	(0.008)	(0.011)	(0.004)	(0.004)	(0.002)
(1) Two-girls	0.027***	0.017***	0.029***	0.028***	0.031***	0.028***	0.026***	0.027***	0.039***	0.027***	0.039***	0.030***	0.028***	0.046***
	(0.004)	(0.007)	(0.002)	(0.006)	(0.006)	(0.010)	(0.006)	(0.008)	(0.002)	(0.008)	(0.012)	(0.004)	(0.004)	(0.002)
(2) More than 2 children	-0.069	-0.102	0.012	-0.265	0.046	-0.118	0.008	-0.037	-0.076	-0.284	0.331	0.060	0.088	-0.056
	(0.112)	(0.455)	(0.060)	(0.169)	(0.152)	(0.207)	(0.169)	(0.290)	(0.049)	(0.270)	(0.293)	(0.115)	(0.098)	(0.040)
Sargan p-value	(0.547)	(0.204)	(0.025)	(0.667)	(0.830)	(0.590)	(0.025)	(0.780)	(0.038)	(0.906)	(0.951)	(0.355)	(0.595)	(0.101)
Observations	138,750	34,515	460,870	52,121	168,170	19,200	52,043	27,423	561,315	27,037	13,043	107,253	106,088	425,546
(3) Parcial-R2	0.0006	0.0002	0.0009	0.0008	0.0012	0.0012	0.0008	0.0006	0.0013	0.0007	0.0010	0.0008	0.0012	0.0026
(4) F-statistic	85.7	6.2	363.7	43.0	63.5	22.5	39.4	16.2	407.7	18.3	13.0	81.0	131.5	835.0

Note: Robust standard errors in parentheses. *** significant at 1%, ** significant at 5%, * significant at 10%. Other covariates in the models are Age, Age at first birth, plus indicators for the sex of the first birth and second children. Samples as described in the text and the appendix. All models include sample weights (see appendix A.3). (1) Coefficient of the first stage using More than 2 children as dependent variable; (2) Coefficient of the second stage using Worked for pay as dependent variable; (3) Goodness of t between More than 2 children and Same-sex after controlling for the other covariates in the model; (4) F-statistic on Same-sex in the first stage regression.

Table A.10: OLS and 2SLS estimates of fertility and female labor supply. 2000s census (MARRIED WOMEN)

	ARG	BOL	BRA	CHL	COL	CRI	ECU	SLV	MEX	NIC	PAN	PER	VEN	USA
Panel A - OLS														
More than 2 children	-0.096***	-0.094***	-0.108***	-0.062***	-0.125***	-0.087***	-0.084***	-0.106***	-0.090***	-0.110***	-0.123***	-0.101***	-0.120***	-0.154***
	(0.003)	(0.006)	(0.002)	(0.004)	(0.005)	(0.007)	(0.004)	(0.007)	(0.002)	(0.007)	(0.009)	(0.003)	(0.003)	(0.002)
Panel B - Instrumental Variables														
(1) Same-sex	0.032***	0.017***	0.031***	0.029***	0.038***	0.034***	0.025***	0.024***	0.034***	0.031***	0.037***	0.029***	0.034***	0.056***
	(0.003)	(0.005)	(0.002)	(0.004)	(0.005)	(0.007)	(0.004)	(0.006)	(0.002)	(0.006)	(0.009)	(0.003)	(0.003)	(0.002)
(2) More than 2 children	-0.223**	-0.026	0.006	-0.211	-0.080	0.024	-0.025	0.022	-0.096**	-0.393**	-0.055	0.016	0.033	-0.073**
	(0.096)	(0.342)	(0.052)	(0.142)	(0.129)	(0.182)	(0.157)	(0.265)	(0.044)	(0.196)	(0.215)	(0.098)	(0.086)	(0.036)
Panel C - Instrumental Variables – over identified														
(1) Two-boys	0.025***	0.009	0.024***	0.022***	0.029***	0.038***	0.019***	0.010	0.023***	0.021***	0.025**	0.019***	0.032***	0.046***
	(0.004)	(0.007)	(0.002)	(0.006)	(0.006)	(0.010)	(0.006)	(0.009)	(0.002)	(0.008)	(0.012)	(0.004)	(0.004)	(0.003)
(1) Two-girls	0.035***	0.021***	0.032***	0.030***	0.039***	0.026***	0.025***	0.028***	0.042***	0.035***	0.043***	0.034***	0.030***	0.056***
	(0.005)	(0.007)	(0.002)	(0.006)	(0.007)	(0.010)	(0.006)	(0.009)	(0.002)	(0.009)	(0.013)	(0.004)	(0.004)	(0.003)
(2) More than 2 children	-0.219***	-0.103	0.020	-0.233	-0.113	0.009	0.039	0.016	-0.082*	-0.412*	-0.045	0.041	0.041	-0.068*
	(0.103)	(0.360)	(0.057)	(0.158)	(0.142)	(0.188)	(0.174)	(0.302)	(0.046)	(0.215)	(0.227)	(0.104)	(0.095)	(0.039)
Sargan p-value	(0.158)	(0.477)	(0.042)	(0.989)	(0.057)	(0.773)	(0.042)	(0.793)	(0.055)	(0.914)	(0.734)	(0.620)	(0.743)	(0.178)
Observations	82,246	31,207	400,605	44,028	139,063	16,422	46,527	21,636	517,595	22,315	11,157	96,117	88,179	312,210
(3) Parcial-R2	0.0013	0.0003	0.0011	0.0010	0.0015	0.0013	0.0008	0.0007	0.0015	0.0012	0.0016	0.0009	0.0014	0.0035
(4) F-statistic	103.4	9.7	371.5	43.9	63.6	22.1	35.1	14.5	430.3	26.4	17.5	90.8	121.9	843.5

*Note: Robust standard errors in parentheses. *** significant at 1%, ** significant at 5%, * significant at 10%. Other covariates in the models are Age, Age at first birth, plus indicators for the sex of the first and second children. Samples as described in the text and the appendix. All models include sample weights (see appendix A.3). (1) Coefficient of the first stage using More than 2 children as dependent variable; (2) Coefficient of the second stage using Worked for pay as dependent variable; (3) Goodness of t between More than 2 children and Same-sex after controlling for the other covariates in the model; (4) F-statistic on Same-sex in the first stage regression.*

Table A.11: Differences in means for demographic variables by Same-sex. 2000 Census

	Arg	Bol	Bra	Chl	Col	Cri	Ecu	Slv	Mex	Nic	Pan	Per	Ury	Ven	USA
Age	-0.079***	-0.044	0.031***	-0.057*	0.012	-0.023	0.021	0.020	0.012	0.009	-0.040	-0.015	-0.066	0.015	0.002
	(0.021)	(0.041)	(0.011)	(0.030)	(0.019)	(0.055)	(0.034)	(0.045)	(0.010)	(0.048)	(0.068)	(0.023)	(0.072)	(0.024)	(0.011)
Age at first birth	0.003	0.029	0.011	-0.022	0.011	0.012	-0.018	0.040	0.008	0.042	-0.034	0.002	-0.065	0.036*	0.002
	(0.018)	(0.033)	(0.009)	(0.028)	(0.015)	(0.045)	(0.027)	(0.038)	(0.008)	(0.034)	(0.057)	(0.019)	(0.066)	(0.019)	(0.011)
Owner	-0.002	-0.007	0.002	0.010***	0.002	0.001	-0.001	-0.004	-0.001	0.002	-0.004	0.001	0.005	-0.003	0.002
	(0.002)	(0.005)	(0.001)	(0.004)	(0.002)	(0.007)	(0.004)	(0.005)	(0.001)	(0.004)	(0.007)	(0.003)	(0.010)	(0.003)	(0.002)
Some primary education	0.001	0.009*	-0.002	0.002	-0.003	0.005	0.001	-0.007	0.000	0.005	-0.002	0.001	0.000	-0.001	-0.000
	(0.002)	(0.005)	(0.001)	(0.003)	(0.002)	(0.006)	(0.004)	(0.006)	(0.001)	(0.006)	(0.007)	(0.003)	(0.007)	(0.002)	(0.000)
Some secondary education	-0.001	-0.012**	-0.001	0.002	0.002	-0.005	-0.001	0.003	0.000	-0.005	0.000	-0.002	-0.003	-0.002	-0.001
	(0.003)	(0.005)	(0.001)	(0.004)	(0.002)	(0.007)	(0.004)	(0.006)	(0.001)	(0.006)	(0.009)	(0.003)	(0.008)	(0.003)	(0.001)
Some tertiary education	0.000	0.004	0.002**	-0.004	0.004	-0.001	-0.000	0.004	-0.001	-0.000	0.002	0.001	0.003	0.003	0.001
	(0.002)	(0.004)	(0.001)	(0.004)	(0.002)	(0.006)	(0.004)	(0.005)	(0.001)	(0.004)	(0.008)	(0.003)	(0.006)	(0.003)	(0.001)
Observations	138,321	34,436	459,385	52,009	167,541	19,137	51,887	27,302	559,994	26,944	13,016	106,982	10,410	105,690	423,549

Source: own estimates based on IPUMS-International. Note: differences in means (mean of Same-sex mothers minus mean of Mixed-sex mothers) and their standard deviations (in parentheses). *** significant at 1%, ** significant at 5%, * significant at 10%. The samples correspond to women aged 21-35 with two or more children aged 18 or younger. The results from Uruguay correspond to 1990 census. Owner is = 1 if the mother owns the house where she lives; Some primary education is = 1 if the maximum education level of the mother is less than complete primary; Some secondary education = 1 for complete primary or incomplete secondary; Some tertiary education = 1 for complete secondary or more.