Fertility and Female Labor Supply in Latin America: New Causal Evidence

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Abstract
In this paper, we study the effect of fertility on maternal labor supply in Argentina and Mexico exploiting a source of exogenous variability in family size first introduced by Angrist and Evans (1998) for the United States. We find that the estimates for the US can be generalized both qualitatively and quantitatively to the populations of two developing countries where, compared to the US, fertility is known to be higher, female education levels are much lower and there are fewer formal facilities for childcare.

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1. Introduction

In Latin America, female labor supply has been increasing since the early 1970's, while at the same time fertility declined. Figure 1 presents these trends for both Argentina and Mexico. The timing of these events suggests that they might be causally related, and one might wonder if the decline in fertility is the cause of the sharp increase in female labor force in these two countries. Despite the relevance of this question, this hypothesis has not been rigorously tested yet in Latin America.

[Insert Figure 1 about here]

One of the most significant changes in human behavior during the past century was the massive incorporation of women into the labor force. Not surprisingly, there is an extensive theoretical and empirical literature attempting to explain female labor supply and its evolution (see, among others, Killingsworth and Heckman, 1986). In particular, the relationship between fertility and female labor supply is of longstanding interest in the social sciences. Much of the research effort has been devoted to disentangling the causal mechanisms linking childbearing and female labor supply (see, among others, Willis, 1987). Recently, Angrist and Evans (1998) (henceforth AE) have made substantial progress in this area. Their identification strategy exploits parental preferences for a mixed sibling sex composition. Parents of same-sex siblings are significantly more likely to have an additional child. Since 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. AE's Instrumental Variables (IV) estimates convincingly show that additional children lead to a reduction in female labor supply.

In this paper, we exploit the AE's identification strategy to estimate the causal effect of childbearing on maternal labor supply in two middle-income Latin American countries: Argentina and Mexico. Using these parameter estimates, we can address whether the decline in fertility observed in the last three decades caused the rise in female labor force participation.

Any causal study faces two sources of threats to its validity: internal and external (see Campbell, 1957, Cook and Campbell, 1979 and Meyer, 1995). Most research is devoted to dealing with threats to internal validity. This refers to whether one can validly draw the inference that, within the context of the study, the differences in the dependent variables were caused by the differences in the relevant explanatory variables. External validity, instead, is concerned with the extent to which a causal relationship holds over variations in persons, settings and time. Although we generally presume that causal relationships can be generalized outside the samples studied, some in our profession have expressed their concern about this. Thus, whenever it is possible, once an identification strategy for a causal relationship is validated internally, it is worth inquiring about the external validity of its results. Ultimately, the external validity of causal estimates is established by replication in other datasets (Angrist, 2004).

This paper also investigates the extent to which the causal link identified in AE can be generalized to the context of developing countries where, compared to the US, fertility is known to be higher, female education levels are much lower and there are fewer formal facilities for childcare. Thus, we question whether in such different socioeconomic environments childbearing also leads to a reduction in female labor supply, and whether the effects are of the same order of magnitude.

The rest of the paper is organized as follows. In the next section we describe the data and provide a set of summary statistics, and we then present the estimation strategy. This is followed by the main estimates of the paper, and a discussion of the exclusion restriction and the implications of the results. Finally, we present a series of robustness checks of the main results. Conclusions follow.

2. Data and summary statistics

Our datasets are gathered from the extended questionnaires samples of both the Mexico 2000 and the Argentina 1991 censuses, conducted respectively by the National Institute of Statistics, Geography and Computing (*Instituto Nacional de Estadística, Geografía e Informática*, INEGI) and the National Institute of Statistics and Censuses (*Instituto Nacional de Estadísticas y Censos*, INDEC). The two result in large and nationally representative datasets. For Argentina, we have data on 16,023,180 individuals and 4,287,580 households, covering around 50 percent of the whole population. For Mexico the sample consists of 10,099,182 individuals and 2,312,034 households, covering around 10 percent of the total population. We restrict our sample to women between 21 and 35 years old, with at least two children, and whose oldest child was at most 18 years old at the time of the census. Following AE, we also exclude from the analysis women whose second child is younger than a year old, and carry out our analysis separately on all women and married women. Thus, our final samples sizes are 599,941 (total) and 456,437 (married) observations for Argentina, and 458,849 (total) and 355,730 (married) for Mexico.

[Insert Table I about here]

Table I presents descriptive statistics and variable definitions. Female employment for our married samples are much higher (30.5 percent) in Argentina than in Mexico (22 percent), but they are both significantly lower than the US figures for equivalent samples (52.8 percent in 1980 and 66.7 in 1990). Both in Argentina and Mexico, female labor supply is lower for married women than for unmarried women. With respect to fertility, the average number of children is higher for married women in Mexico (3.035) than in Argentina (2.985), and substantially higher than the respective US figure (around 2.5 in both the 1980 and 1990 censuses).

In this paper, the fertility variable of interest – i.e., the causing variable in our empirical labor supply regression models – is the indicator *More than two children*, which is instrumented by the indicators: *Same sex*, *Two boys* and *Two girls*. In both Argentina and Mexico, slightly above 50 percent of the women in any of the samples considered have a third child while in the US the same figure is only about 36 to 40 percent. We also report indicators for whether the first and second children were boys. Finally, the table also presents the women's age and age at first birth.

3. Estimation strategy

Let D_i be an indicator for women with more than two children in a sample of women with at least two children. Additionally, let Y_{1i} be the labor supply of mother i if D_i equals 1 and Y_{0i} denote her labor supply otherwise.

The simplest option to estimate the causal effect of childbearing on labor supply is by means of the following linear, constant-effects model (Angrist, 2001):

$$E[Y_{0i}] = X_i' \beta \qquad (1a)$$

and

$$Y_{1i} = Y_{0i} + \alpha \qquad (1b)$$

where X is a vector of control variables. These assumptions lead to the following linear causal model:

$$Y_i = X_i'\beta + \alpha D_i + \varepsilon_i \qquad (2)$$

which is easily estimated by Two-Stages Least Squares (2SLS) if the IV Z_i (*Same Sex*) satisfies: a) $\{Y_{0i}, Y_{1i}, D_{0i}, D_{1i} \mid X_i\}$ is independent of Z_i, b) $P[D_i = 1 \mid X_i, Z_i] \neq P[D_i = 0 \mid X_i, Z_i]$, and c) without loss of generality, $D_{1i} \geq D_{0i}$.

In general, if the additive, constant-effect assumptions in (1a) and (1b) do not hold, a 2SLS estimate of model (2) does not identify the average causal effect without further assumptions, but identifies a Local Average Causal Effect (LATE), i.e. the average effect of treatment on those individuals whose treatment status is induced to change by the instrument (see Imbens and Angrist, 1994). Angrist *et al.* (1996) refer to this group as the population of compliers.

We present estimates of the effect of childbearing on labor supply by Ordinary Least Squares (OLS) and 2SLS. As argued in Angrist (2001), the problem of causal inference with Limited Dependent Variables (LDV) is not fundamentally different from causal inference with continuous outcomes. If there are no covariates or the covariates are sparse and discrete, linear models and associated estimation techniques like 2SLS are no less appropriate for LDV than for other types of dependent variables. This is because conditional expectation functions of discrete covariates can always be parameterized as linear in the parameters by saturating the model, regardless of the support of the dependent variable. Thus, we first present 2SLS estimates of the parameter of interest from models where we saturate the whole set of control variables. However, since the potential outcome conditional expectation function (CEF) is also a function of the causing variable, we denote this empirical model as an "almost saturated" model. We then present 2SLS estimates of the parameter of interest from more parsimonious models. Finally, as a robustness check, we also report estimates from the IV estimator developed by Abadie (2003), which allows a flexible nonlinear approximation of the causal response function. Abadie's (2003) "Causal

IV" estimates have a robust causal interpretation regardless of the shape of the actual CEF for potential outcomes, since identification is attained non-parametrically.

4. Main Results

4.1. Wald Estimates

Table II contains the essence of the estimation strategy as well as our main results. First, note that in both Argentina and Mexico the probability of having two children of the same sex is just above one-half. As expected, in both countries women who have had two children of the same sex have a higher probability of having a third child (and, naturally, also a higher number of children) than women who have had two children of different sex (*Mixed sex*). The difference in these probabilities is around 3.5-4 percentage points in Argentina (all-married) and 3.2-3.6 percentage points in Mexico (all-married); women with two children of the same sex have on average 0.06-0.07 more children in both countries. These differences, significant at the 1 percent level, are close to those found by AE for the US (0.07-0.08 more children), and represent evidence of a sex mix preference phenomenon in both Argentina and Mexico.

With respect to labor supply, there is also evidence of a small but significant difference in employment between women who have had two children of the same sex and women with mixed sex siblings. The latter group has a participation rate around 0.3-0.4/0.2-0.3 percentage points higher (Argentina/Mexico), and these differences are statistically significant at the standard levels of confidence. In the US, AE find a similar pattern, although the differences are larger (around 0.8 percentage points for 1980 and 0.5 for 1990).

The last four lines in the two panels report Wald and OLS estimates of the effect of the total number of children, and having more than two children, on the probability of working for pay. Wald estimates are obtained straightforwardly as the ratio of the reduced-form relationship between *Worked for pay* and *Same sex*, and between *Number of children* or *More than two children* and *Same sex*. All Wald estimates are statistically significant at conventional levels of confidence. The Wald estimates are equivalent to the simplest instrumental variable estimates obtained by relying on *Same sex* as an instrument for *Number of children* or *More than two children*, when no other covariates are included in the CEF.

[Insert Table II about here]

The results for Argentina imply that an additional child reduces the labor supply of women whose fertility has been affected by their children's sex mix by about 5-6 percentage points, while having more than two children reduces their labor supply by about 9-10 percentage points. For Mexico, the effects are quite similar: 3.6-4.9 and 6.8-9.2 percentage points respectively. These results are quite close to the 1990 US estimates reported by AE (6.3 and 8.4 percentage points respectively), although they are lower than the US 1980 effects (13.3 and 10.4 percentage points respectively). Finally, it should be noted that both LATE estimates are larger than the simple OLS estimates (in absolute value). It is worth remembering that, in general, the *Same sex* IV strategy identifies the average effect of having more than two children on those whose fertility decisions are changed by the instrument (compliers), while OLS is suspected to fail at identifying the same effect averaged over the whole population. Thus, with this interpretation in mind, the finding is not worrisome.[1]

4.2. Exclusion restriction

Exclusion restrictions are non-testable directly, and their plausibility must be evaluated on a case-by-case basis. In some developing countries, there might be concerns that the presence of strong son preferences could affect the sex composition of children, either through stopping rules or selective abortion, violating the exclusion restriction. However, Figure 2 rules out this concern in our samples. As can be seen, both in Argentina and Mexico, and in the US, the infant sex ratios (the ratio of boys to girls aged zero to four) are approximately equal to the biological ones, which are about 1.04. On the contrary, in China and Korea, the infant sex ratios show evidence of parental actions affecting biological sex ratios (see, among others, Basu *et al.*, 2003).

[Insert Figure 2 about here]

Basu and Das Gupta (2001) also argue that beyond cultural and religious factors, some societies exhibit a strong son preference because of the gap between sons' and daughters' "ability to contribute to the physical, emotional and financial well-being of their parental

[1]See Card (2000) for a similar argument in reconciling IV and OLS estimates of the effect of schooling on earnings.

household." This gap is mainly determined by kinship systems: if women's links with their parents are cut-off after marriage, it becomes more attractive to rear sons that will be able to provide for their parents in old age. Similarly, in some societies where parents are expected to pay substantial dowries for their daughters, rearing girls is relatively more expensive than rearing boys. In cases like these, the permanent income of parents is affected by the sex mix of children, which may in turn have an effect on their labor supply.

Family institutions both in Argentina and Mexico, however, do not exhibit any severe form of son preference. Dowries are virtually unheard of in both countries, and extreme preferences for sons imply forms of discrimination against girls that are not observed in Latin America in general. For instance, Figure 2 depicts the lack of a systematic effect of son preference on the mortality of girls, and Table III shows that both in Argentina and Mexico the primary school enrollment rates are virtually the same for boys and girls – if anything, the average differences appear to be smaller than in the United States.

[Insert Table III about here]

Finally, another possible threat to the validity of the identification strategy is posed by Rosenzweig and Wolpin (2000). Studying outlays per children in rural India, they find that same sex siblings are related to substantially lower levels of expenditure. They attribute this effect to "hand-me-down" savings, which are more likely to arise when there are children of the same sex in the household for items such as clothing and footware. Since these items represent a sizeable fraction of the household's expenditures, they note that the sex composition of children plausibly alters labor supply through mechanisms other than through fertility change alone.

While expenditure data per child is not available for Argentina or Mexico, survey data suggests that sex composition is unlikely to have a noticeable effect on expenditure. Rosenzweig and Wolpin (2000) find in their Indian data that clothing expenditures on children under 18 represents 11 percent of household income. For Mexico, Hernández Franco and Pérez García (2003) report that in the year 2000 households spent around 4.8 percent of their budget on clothing and footwear for all members of the family, with little variation among deciles of household income. Meanwhile, Argentine households in 1987 devoted 6.7 percent of their budget to the same items (for all members), and only 2.8 percent on clothing and foot wear for children aged 10 or less.[2]

Rosenzweig and Wolpin's (2000) estimated "hand-me-down" savings for these goods amounts to 1.3 percent of average earnings: even assuming that these savings exist in Argentina and Mexico (and that they imply a direct effect on labor supply), their size would be too small to account for a meaningful reduced form relationship between a same sex indicator and parental labor supply.

Thus, the evidence presented in this section supports the exogeneity of the *Same sex* indicator and the internal validity of the Wald estimates presented above.

4.3. Discussion

The basic AE results are thus also qualitatively valid in the two developing countries considered. The next question is whether the effects are of the same order of magnitude or whether they differ substantially. A test of the hypotheses that the effect of fertility on female employment for Argentina (1991), Mexico (2000) and the United States (1990, from AE) does not reject the null at standard levels of statistical significance. Thus, we can assert that in the US, and in Argentina and Mexico, the average effect of going from a family size of two children to more than two is statistically similar (for those whose treatment status is changed by the *Same sex* instrument). Thus, AE estimates appear to be generalized both qualitatively and quantitatively to dissimilar populations that display obvious observable differences.

Both in Argentina and Mexico, childbearing also leads to a reduction in female labor supply – but how much of the increase in labor force participation in the last three decades can be explained by the large changes in fertility observed in Figure 1? According to CELADE (2004), fertility fell by 14 percent in Argentina and almost 60 percent in Mexico between 1970 and 2000, while female labor force participation increased by 11.8 percentage points in Argentina (almost 60 percent) and by 15.9 percentage points in Mexico (an increase of 153 percent). Using the Wald estimates for the complete samples in Table II, one fewer child implies a fall in female labor force participation of about 5-3.5 percentage points

[2] The figures for Argentina are based on the 1987 Expenditure Survey by INDEC. Based on further evidence from this survey (available upon request), we failed to find any effect of the sex composition of children on the budget shares of clothing, education, food and other categories of goods.

(Argentina-Mexico). The modest fall of 0.4 children in Argentina between 1970 and 2000 accounts for an increase of only 2.1 percentage points in participation, which in turn can only explain 18 percent of the increase in female employment during the same period. In Mexico, however, the very large reduction in fertility from 6.8 to 2.8 children accounts for an increase of 14.5 percentage points in participation, which is most of the 15.9 points rise during this period.

5. Robustness Checks

The estimates in the previous section correspond to a simple version of model (2) where only the causing variable is included in the estimation. In this section we check their robustness. First, we include a set of standard control variables: age of the woman, her age at first birth, sex of the first child and sex of the second child (see Angrist and Evans, 1998). In order to saturate the model, we map both age and age at first birth into five categories each, then create a set of forty-nine mutually exclusively indicators by interacting them with the aforementioned control variables.[3,4]

Conditioning on the sex of the first two children allows us to control for any secular additive effect of child gender on female participation. This is useful because *Same sex* is potentially correlated with the sex of either child, which is of concern if this affects labor supply for reasons other than family size.

Table IV presents these estimates, which are almost identical to those presented in the previous section.[5] A third child reduces the probability of work by about 6.31-8.62 percentage points (all-married) for Mexico and 8.17-9.58 percentage points (all-married) for Argentina. Most of these coefficients are significantly different from zero at the 5 percent level (the exception being 10 percent for the total sample in Mexico).[6]

[3] The five age category indicators were chosen to contain approximately the same number of observations in each of them, and were defined as 21-25, 26-28, 29-30, 31-32 and 33-35 for age, and 17 or less, 18-19, 20-21, 22-23 and 24 and more for age at first birth.

[4] We also fitted a more parsimonious version of these models including controls for the continuous variables *Age* and *Age at first birth*, and indicators for the sex of the first and second child, instead of interactions of categorized versions of these variables. The results were identical, and are available upon request as a separate appendix.

[5] This result conforms with our identification strategy, since it shows that the instrument is orthogonal to the additional covariates.

[6] These results are also robust to: a) using labor for participation as the dependent variable instead the alternative *Worked for pay*; b) including in the sample women whose second child is younger than a year old; and

In Table IV we also present Abadie's (2003) IV estimates.[7] The results are almost identical to the 2SLS just described, showing that our almost saturated model captures extremely well the CEF of female labor supply. This finding coincides with those presented in Angrist (2001) and Abadie (2003).

[Insert Table IV about here]

In addition, the *Same sex* indicator is easily decomposed into two variables indicating the sex composition of the first two children, *Two boys* and *Two girls*, leading to an overidentified model. AE show that this is useful because the bias from any secular effects of child gender on labor supply should be different from these two instruments, while the labor supply consequences of childbearing seem likely to be independent of whether *Same sex* equals *Two boys* or *Two girls*. Thus, an appropriate specification test is the Sargan test or test of overidentifying restrictions. However, when *More than two children* is instrumented by both *Two boys* and *Two girls*, it is not possible to control for the effects for the sex of each child, and so we report results that control only for the sex of the first child (as in AE).

Table IV shows that women who have two girls have on average a 4.7-5.3 (all-married, Argentina) and 4.3-4.6 (all-married, Mexico) percentage point higher probability of having a third child, while these figures are lower for women who have had two boys (around 2.6-3 and 2.5-2.8 percentage points, respectively). Nevertheless, the *Two boys* variable indicate the presence of mixed sex sibling preference in both countries.[8]

These first-stage results on sex preferences are qualitatively similar to what AE report for the United States. As in our data, they find a difference in the probability of further childbearing between women who have had two boys, and women who have had two girls in their 1980 dataset (although not in their 1990 dataset).

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c) including in the sample some women that were discarded because of mismatches (see data appendix). We also added municipality dummy variables and, again, the estimates did not change at all. Finally, the results were unaltered when we included the spouse controls instead of the women's. All of these results are available upon request as a separate appendix.

[7] We implement a simple two-step version of Abadie's (2003) estimator based on a linear specification of the Local Average Response Function. In the first step, we estimate by OLS the model $Z_i = X_i'\gamma + e_i$. The predicted values from this regression are then used to construct the estimated weights K_i (Abadie 2003, Theorem 3.1). We finally use these weights to estimate the model given by equation (2), $Y_i = X_i'\beta + \alpha D_i + \varepsilon_i$, by weighted least squares (Abadie 2003, Equation 14).

[8] Strict son preference with no mixed sex sibling requires a coefficient of *Two boys* not significantly different from zero (Leung, 1991).

The Generalized Instrumental Variables Estimates (GIVE) are smaller (in absolute terms) than the IV ones. However, the differences are never statistically significant at conventional levels of significance. In addition, the statistics of contrast of a standard Sargan test of over-identifying restrictions are reassuring. We cannot reject the null hypothesis at the 5 percent level for any of the four samples considered.

6. Conclusion

We study the effect of women's fertility on labor supply in Argentina and Mexico exploiting an instrumental variable estimator first introduced by Angrist and Evans (1998) for the United States. Our investigation shows that the "mixed sex sibling preference", the basis of AE identification strategy, is present in Argentina and Mexico. Moreover, we provide new supporting evidence of the internal validity of the AE identification strategy of the causal effect of fertility on labor market outcomes for families with at least two children. More importantly, we find that the AE estimates for the US can be generalized both qualitatively and quantitatively to the populations of two developing countries which are notably different from the original application.

Both in Argentina and Mexico, childbearing also leads to a reduction in female labor supply. Our estimates suggest that the decline in fertility in Argentina can only account for 18 percent of the rise in female labor force participation during the last three decades. Much to the contrary, the very large drop in fertility observed in Mexico during the same period accounts for most of the increase in female employment.

Appendix: Data sources

The Argentine dataset contains information on 16,023,180 individuals and 4,287,580 households, from a total population of 32,245,467 individuals and 8,927,289 households. We constructed this dataset from the original data tapes. The Mexican dataset covers 10.6 percent of the total population of 97,483,412 persons and 22,268,916 households, yielding a sample size of 10,099,182 individual records and 2,312,034 household records. The Mexican data is taken from Sobek *et al.* (2002).

Matching women and their children

For both Argentina and Mexico, the relationship variable linking members of a household indicates kinship with respect to the head of the household only. In order to match women with their own children, we restrain the sample to females who are heads or spouses of the heads of households. In order to avoid assigning all children of a male head to his current partner, who might not be the mother of all children in the household, we first check that the reported number of children alive (as asked for in a specific census question for both countries) coincides with the number of children in the household matched to the woman, restraining our samples to women for whom both numbers coincide. Finally, we made some extra adjustments based on the age of the woman and/or her husband. We discard a small number of observations for which the age of the mother at first birth was less than 14, taking this as an indicator of data entry errors or misallocated children, since most of the ages were far too low. We also dropped from our final samples a very small fraction of married women for whom the husband's age at first birth was less than 14.

Worked for pay indicator

In the Argentine census, an individual is classified as working for pay (*Worked for pay* indicator equal to 1) if he or she works and is not a family worker without remuneration. Thus individuals working for pay include employees (wage earners), the self-employed, owner-managers and civil and domestic servants. In Mexico, we use the same definition and classify an individual as working for pay if he or she does remunerated work.

References

Abadie, A. (2003): "Semiparametric Instrumental Variable Estimation of Treatment Response Models", *Journal of Econometrics* 113, pp. 231-63.

Angrist, J. (2004): "Treatment Effect Heterogeneity in Theory and Practice", *The Economic Journal*, 114 (494), pp. C52-C83.

Angrist, J. (2001): "Estimation of Limited Dependent Variable Models with Dummy Endogenous Regressors: Simple Strategies for Empirical Practice", *Journal of Business and Economic Statistics* 19, pp. 2-16.

Angrist, J., G. Imbens and D. Rubin (1996): "Identification of Causal Effects Using Instrumental Variables", *Journal of the American Statistical Association* 91, pp. 444-455.

Angrist, J. and W. Evans (1998): "Children and their Parents' Labor Supply: Evidence from Exogenous Variation in Family Size", *American Economic Review* 88(3), pp.450-77.

Basu, A. and Das Gupta, M. (2001): "Family Systems and the Preferred Sex of Children." In Hoem, J. editor, (2001), *International Encyclopedia of Social and Behavioral Sciences*, Elsevier.

Card, D. (2000): "The Causal Effect of Education on Earnings" in Orley Ashenfelter and David Card (eds.), *Handbook of Labor Economics*, Vol. 3A. Amsterdam: North-Holland.

Campbell, D. T. (1957): "Factors Relevant for the Validity of Experiments in Social Settings", *Psychological Bulletin* 54, pp. 297-312.

CELADE (2004), "Population Estimations and Projections for Latin America and the Caribbean", *Demographic Bulletin*, No. 79.

Cook, T. and D. T. Campbell (1979): *Quasi-Experimentation: Design and Analysis issues for field settings*. Boston: Houghton Mifflin.

Das Gupta, M., Zhenghua, J., Zhenming, X., Bohua, L., Chung, W. and Hwa-Ok, B., (2003): "Why is Son Preference so Persistent in East and South Asia? A cross-country study of China, India, and the Republic of Korea", *Journal of Development Studies* 40(2), pp. 153-187.

Hernández Franco, D. and Pérez García, M. (2003): "En el año 2000. Gasto de los Hogares y Pobreza en México", *Cuadernos de Desarrollo Humano*, No. 5, Secretaría de Desarrollo Social, México.

Imbens, G. and J. Angrist (1994): "Identification and Estimation of Local Average Causal Effects", *Econometrica* 62, pp. 467-75.

Killingsworth, M. and J. Heckman (1986): "Female Labor Supply: A Survey" in Orley Ashenfelter and Richard Layard (eds.), *Handbook of Labor Economics*, Vol. 1. Amsterdam: North-Holland.

Leung, S. F. (1991): 'A Stochastic Dynamic Analysis of Parental Sex Preferences and Fertility", *Quarterly Journal of Economics*, Vol. 106(4), pp. pp. 1063-88.

Meyer, B. (1995): "Natural and Quasi-Experiments in Economics", *Journal of Business & Economic Statistics* 13(2), pp. 151-62.

Pantelides, E. (2002): "Completing the Fertility Transition: The Case of Argentina", in United Nations (2002).

Rosenzweig, M. and K. Wolpin (2000): "Natural 'Natural Experiments' in Economics", *Journal of Economic Literature*, Volume 38(4), pp. 827-874.

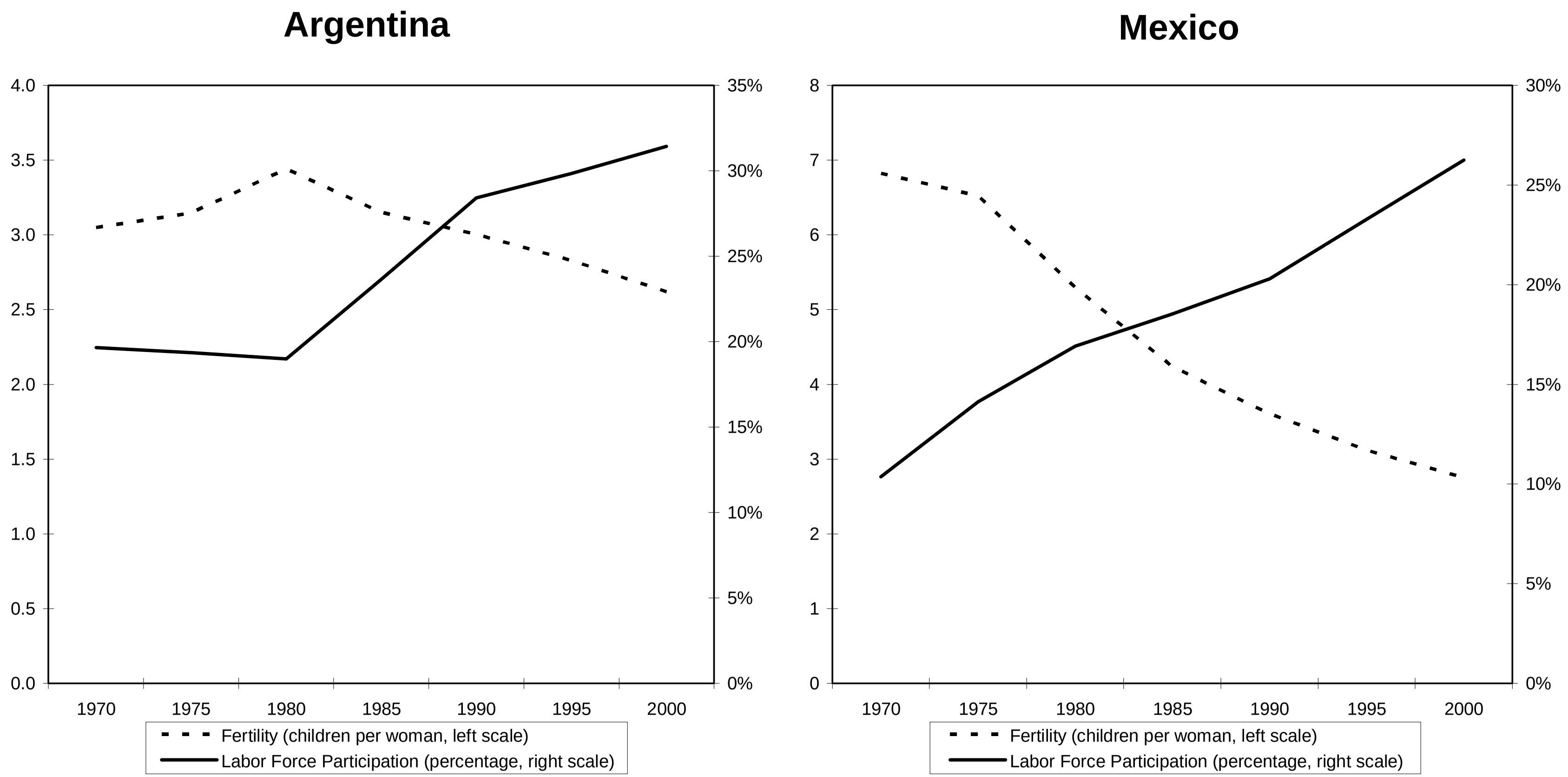
Sobek, M., S. Ruggles and R. McCaa (2002): "Integrated Public Use Microdata Series-International: Preliminary Version 1.1", Minneapolis: Minnesota Population Center, University of Minnesota.

Tuiran, R., Partida, V., Mojarro, O. and Zúñiga, E. (2002): "Fertility in Mexico: Trends and Forecast", in United Nations (2002).

United Nations (2002): *Completing the fertility transition*, Report of the Expert Group Meeting, Department of Economic and Social Affairs, Population Division, United Nations, New York.

Willis, R. (1987): "What Have We Learned from the Economics of the Family?", *American Economic Review* 77(2), pp. 68-81.

Figure 1: Fertility and female labor force participation 1970-2000, Argentina and Mexico



Source: CELADE (2004).

Table I - Summary statistics

	Argentina 1991		Mexico 2000	
	All women	**Married women**	**All women**	**Married women**
Worked for pay	0.315	0.305	0.239	0.220
(=1 if worked for pay, 0 otherwise)	(0.465)	(0.460)	(0.426)	(0.414)
More than 2 children	0.596	0.574	0.592	0.593
(=1 if mother had more than two children, 0 otherwise)	(0.491)	(0.495)	(0.491)	(0.491)
Number of children	3.062	2.985	3.029	3.035
	(1.240)	(1.183)	(1.188)	(1.197)
Same Sex	0.506	0.505	0.503	0.503
(=1 if first two children were the same sex, 0 otherwise)	(0.500)	(0.500)	(0.500)	(0.500)
Two boys	0.260	0.261	0.261	0.261
(=1 if two children were boys, 0 otherwise)	(0.438)	(0.439)	(0.439)	(0.439)
Two Girls	0.246	0.244	0.243	0.242
(=1 if two children were girls, 0 otherwise)	(0.431)	(0.430)	(0.429)	(0.428)
Boy 1st	0.508	0.510	0.512	0.512
(=1 if first child was a boy, 0 otherwise)	(0.500)	(0.500)	(0.500)	(0.500)
Boy 2nd	0.506	0.507	0.507	0.507
(=1 if second child was a boy, 0 otherwise)	(0.500)	(0.500)	(0.500)	(0.500)
Age	29.660	29.928	29.440	29.651
	(3.770)	(3.652)	(3.758)	(3.683)
Age at first birth	20.641	20.932	19.930	20.095
	(3.337)	(3.340)	(3.083)	(3.101)
Observations	599,941	456,437	458,849	355,730

Note: Means and standard deviations (in parentheses). The samples correspond to the extended questionnaire sample of the 1991 Census, Argentina and the 2000 Census, Mexico. Samples as described in the data appendix.

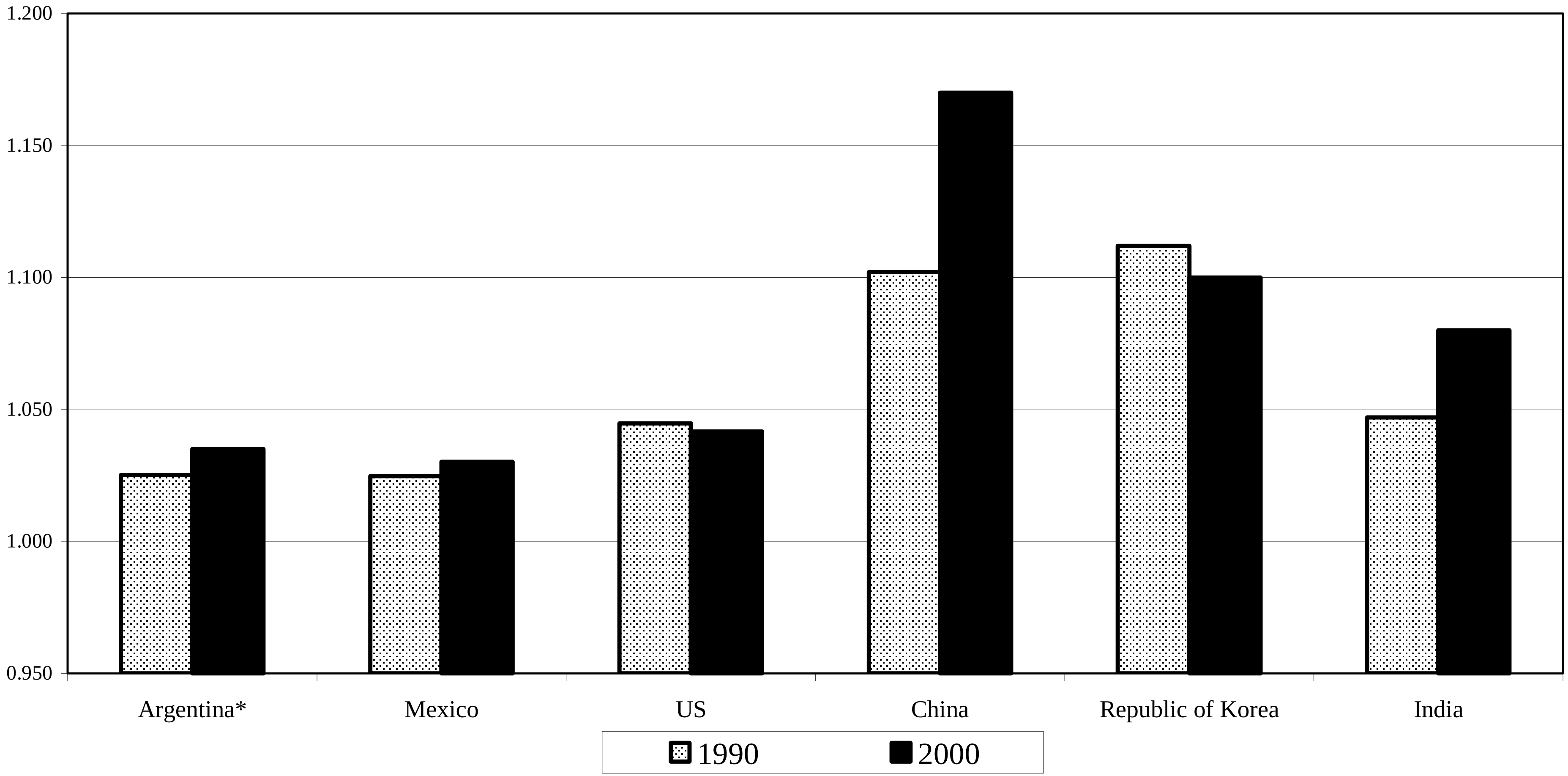
Table II - Wald estimates

	Argentina							
	All women				Married women			
	Proportion of sample	Worked for pay	Number of children	More than two children	Proportion of sample	Worked for pay	Number of children	More than two children
Total		0.3155	3.0619	0.5955		0.3046	2.9848	0.5741
Same sex (1)	0.5062	0.3139	3.0935	0.6131	0.5053	0.3025	3.0196	0.5940
Mixed sex (2)	0.4938	0.3171	3.0295	0.5775	0.4947	0.3066	2.9492	0.5538
Difference (1)-(2)		-0.0032	0.0640	0.0356		-0.0041	0.0704	0.0403
Wald estimate			-0.0503	-0.0905			-0.0584	-0.1021
Standard error			[0.0187]***	[0.0336]***			[0.0193]***	[0.0337]***
OLS estimate			-0.0383	-0.0913			-0.0410	-0.0875
Standard error			[0.0005]***	[0.0012]***			[0.0006]***	[0.0014]***
Observations			599,941				456,437	

	Mexico							
	All women				Married women			
	Proportion of sample	Worked for pay	Number of children	More than two children	Proportion of sample	Worked for pay	Number of children	More than two children
Total		0.3155	3.0619	0.5955		0.3046	2.9848	0.5741
Same sex (1)	0.5034	0.2378	3.0598	0.6080	0.5030	0.2180	3.0684	0.6108
Mixed sex (2)	0.4966	0.2400	2.9981	0.5754	0.4970	0.2213	3.0010	0.5747
Difference (1)-(2)		-0.0022	0.0617	0.0326		-0.0033	0.0674	0.0362
Wald estimate			-0.0357	-0.0677			-0.0490	-0.0914
Standard error			[0.0187]***	[0.0385]*			[0.0193]***	[0.0383]**
OLS estimate			-0.0274	-0.0609			-0.0278	-0.0620
Standard error			[0.0203]*	[0.0013]***			[0.0206]**	[0.0014]***
Observations			458,849				355,730	

Note: * significant at 10%; ** significant at 5%; *** significant at 1%. The samples correspond to the extended questionnaire sample of the 1991 Census, Argentina and the 2000 Census, Mexico. Samples exclude women whose second child is less than a year old, as described in the data appendix.

Figure 2: Infant (0-4 years) male to female sex ratios, selected countries, 1990-2000



Source: Argentina, Mexico and United States: authors' calculations based on the respective census and the International Data Base (US Census Bureau). China, Korea and India: Basu and Das Gupta (2003).
*Note: Values for Argentina are from the 1991 and 2001 Census.

Table III - Enrollment rates

	Argentina, 1991			Mexico, 2000			United States, 1990		
Age	**Boys**	**Girls**	**Difference**	**Boys**	**Girls**	**Difference**	**Boys**	**Girls**	**Difference**
5	83.61%	83.84%	0.23%	72.63%	72.79%	0.15%	67.82%	68.36%	0.54%
6	95.64%	95.99%	0.35%	89.83%	89.93%	0.09%	92.23%	92.44%	0.20%
7	96.96%	96.60%	-0.36%	95.19%	95.25%	0.06%	94.61%	94.73%	0.12%
8	97.72%	97.90%	0.17%	96.16%	96.30%	0.15%	95.29%	95.38%	0.10%
9	97.76%	97.87%	0.12%	96.61%	96.71%	0.11%	95.66%	95.83%	0.17%
10	97.37%	97.77%	0.41%	96.15%	96.44%	0.29%	95.46%	95.75%	0.28%
11	97.06%	97.34%	0.27%	96.04%	96.15%	0.10%	96.01%	96.14%	0.13%
12	95.78%	96.12%	0.34%	92.90%	91.85%	-1.05%	96.27%	96.46%	0.19%
Total	95.25%	95.43%	0.19%	91.91%	91.90%	-0.01%	91.55%	91.81%	0.25%

Note: Authors' calculations from the respective Census (INDEC for Argentina and Sobek *et al.* , 2002, for Mexico and the United States). Samples include all children in each age category. The difference is the rate for girls minus the rate for boys.

Table IV - First and second stages, almost saturated model

	Argentina		Mexico	
	All women	Married women	All women	Married women
First stage - dependent variable: More than two children				
Coefficient of:				
Same Sex[1]	0.0366	0.0413	0.0336	0.0371
	[0.0012]***	[0.0014]***	[0.0013]***	[0.0015]***
Two Boys[2]	0.0260	0.0300	0.0247	0.0284
	[0.0017]***	[0.0019]***	[0.0019]***	[0.0021]***
Two Girls[2]	0.0475	0.0529	0.0429	0.0461
	[0.0017]***	[0.0019]***	[0.0019]***	[0.0021]***
Second stage - instrumented variable: More than two children				
OLS[1]	-0.0969	-0.0828	-0.0903	-0.0812
	[0.0013]***	[0.0015]***	[0.0014]***	[0.0015]***
IV: Same Sex[1]	-0.0817	-0.0958	-0.0631	-0.0862
	[0.0323]**	[0.0325]***	[0.0370]*	[0.0370]**
DWH p-value	0.6361	0.6868	0.4628	0.8929
IV: Same Sex[1] - Abadie's estimator	-0.0814	-0.0953	-0.0631	-0.0862
	[0.0333]***	[0.0378]***	[0.03962]*	[0.0415]**
IV: Two Boys and Two Girls[2]	-0.0652	-0.0821	-0.0445	-0.0721
	[0.0310]**	[0.0313]***	[0.0357]	[0.0360]**
Sargan p-value	0.0701	0.1121	0.0545	0.1015
DWH p-value	0.3050	0.9825	0.1994	0.8006
Observations	599,941	456,437	458,849	355,730
Geographic Controls	None		None	

Note: Standard errors in brackets. * significant at 10%; ** significant at 5%; *** significant at 1%. [1]Control for sex of first and second children. [2]Control for sex of first child. All regressions include main effects and interactions for five categories of age, five categories of age at first birth, and sex of the first children (49 indicator variables in total). Standard errors for Abadie's estimator were obtained by 500 bootstrap replications. Samples as described in the data appendix.