Motherhood and Female Labor Force Participation: Evidence from Infertility Shocks*

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Abstract

We propose a new exogenous source of variation in family size based on infertility shocks to find the causal effect of children on female labor force participation. Clearly, infertility affects the number of children a women can have. In addition, other than the fact of increasing with age, infertility is virtually random. Thus, an indicator variable for the infertility status of women of childbearing age is a plausible instrument for childbearing. We apply this methodology to a sample of women in six Latin American countries, where OLS estimates suggest a negative relationship between children and women's labor force participation. Our main finding is that, after using the infertility instrument, there is no evidence that children have a causal effect on the labor force participation of women.

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1 Introduction

Women are underrepresented in the paid labor force in both developed and developing countries. Recent scholarship has argued that labor force differentials are not due to gender per se, but can be attributed to the fact that women disproportionately face the responsibilities associated with bearing and raising children[1]. While the negative relationship between the presence of children and participation in the labor force is well established, interpretation of this relationship is complicated by the endogeneity of fertility. The number of children a woman has is a choice variable which could be influenced by her labor force participation. Additionally, there are likely to be omitted factors that influence both fertility and labor force participation. For instance, women with high career-based unobservables (such as ambition or talent) may choose to have fewer (or no) children and these women may thus be overrepresented in the labor force. Thus the observed negative relationship between children and labor force participation could be spurious.

Several studies have exploited exogenous changes in family size in order to identify the causal relationship between the number of children and female labor supply. Examples of this approach include twins at first birth (e.g. Rosenzweig and Wolpin 1980, Bronars and Grogger 1994, Jacobsen, Pearce III, and Rosenbloom 2001) and the sex composition of the first two children (e.g., Angrist and Evans 1998, Cruces and Galiani 2007).

We propose an alternative exogenous source of variation in family size based on infertility shocks to find the causal effect of children on female labor force participation. Clearly, infertility affects the number of children a women can have. In addition, other than the fact of increasing with age, infertility is virtually random. Thus, an indicator

[1] See Waldfogel (1998) for a general discussion of the family gap and Piras and Ripani (2005) for evidence that a family gap exists in Latin America.

variable for the infertility status of women of childbearing age is a plausible instrument for childbearing[2].

An advantage of this new instrument is that, unlike previous studies that use twinning or sex mix to generate variation in the number of children, our empirical strategy can investigate the differential labor supply between childless women and women with children. Thus we are able to identify the causal effect of having children on female labor force participation for a broader sample of women.

We apply this methodology to a sample of women in six Latin American countries, where OLS estimates suggest a negative relationship between children and women's labor force participation. Our main finding is that, after using the infertility instrument, there is no evidence that children have a causal effect on the labor force participation of women.

2 Background information on infertility

The medical literature defines infertility as the failure to conceive after a year of regular intercourse without contraception. This definition is further broken down into primary infertility, which describes women who have never been able to conceive a pregnancy, and secondary infertility, describing those who have had at least one successful pregnancy, but have not been able to achieve another. It is well established that infertility (both primary and secondary) increases with a woman's age (Dunson et al, 2004). However, the medical literature is not in agreement about what other factors, if any, influence infertility. There is some evidence suggesting that indicators of poor health such as sexually transmitted diseases, high body mass index (BMI), and miscarriages are associated with infertility (Grodstein et al, 1994). However, infertility appears to

[2]Our identification strategy most closely resembles Cristia (2006) who investigates differences in employment outcomes by child status for women seeking help to become pregnant.

be a random event in that the mother's background characteristics are unrelated to observed heterogeneity in fertility (Joffee and Barnes, 2000).

3 Demographic and health surveys

In this paper we use cross-sectional data from the Demographic and Health Surveys (DHS) in Peru (conducted in 1996), Guatemala (1998), Colombia (1995), Bolivia (1994 and 1998), Nicaragua (1998) and the Dominican Republic (1996). The DHS are standardized nationally representative household surveys in developing countries. Women answered questions about their employment status, birth history, current and future contraceptive use, fertility preferences and their socio-economic, marital and health status.

We identify self-reported infertility in two ways. The first is when women mentioned sub-fertility or infertility as their reason for not currently using contraceptives. In the second, when asked about their desire for future children, non-sterilized women responded that they are unable to have more children. We define a woman as infertile in either of these cases. In keeping with the medical definition of infertility, we can only identify infertility for non-sterilized women who are not currently taking contraceptives. These women constitute more than 60% of the sample. We exclude from the sample students, women who were using contraceptives, sterilized women and women have never had a sexual encounter. Our main sample contains 24,131 women between the ages of 20 and 44. When including health indicators, asked to only a sub-sample, the number of observations is reduced to 15,992.

Our labor force participation variable takes the value of one if a woman reported working for pay during the week prior to the survey and zero otherwise. We define the number of children in three separate ways: the number of children living at home, the number of children under the age of six, and a binary variable equal to one if the woman has at least one child and zero otherwise. In the sample, 52% of women participated in the labor force, 7.3% report being infertile, 84% have at least one child and the average woman has 2.3 children living at home including one child under the age of six.

4 Methodology

For the sample described above, the main specification is given by:

$$LFP_i = \alpha + \beta K_i + \sum_j \gamma_j \text{AGE}_{ji} + X_i'\delta + e_i \tag{1}$$

where LFP_i is equal to 1 if the i-th woman is in the labor force and zero otherwise. The key variable is K_i and it captures the number/presence of children living at home. Thus, β is the parameter of interest. Because of the nature of infertility we include the woman's age in the form of binary age-group categories (indexed by j) in all specifications. Vector X_i varies by model. Model 1, the most parsimonious model, includes categorical indicators of educational attainment, age, age and education interactions and country fixed effects. Model 2 contains Model 1 and adds control variables that may influence labor force participation such as age at first intercourse, marital status, age at first marriage, and spouse's education. Finally, Model 3 contains all of the variables in Model 2 plus an indicator of health status.

OLS estimates of β are likely to be biased due to unobserved variables in e_i. The direction of bias is given by two elements: the relationship between the omitted variable and the outcome variable (LFPi) and its relationship with the variable of interest (K_i). In particular, consider the case when a driving force for women to join the labor market is their career ambition. If ambition correlates positively with the outcome variable and negatively with the number of children, excluding this variable from equation (1) biases the OLS estimates upwards since part of the estimated effect of children on labor force participation is actually due to ambition. We will use infertility to instrument for K_i in equation (1) to address the endogeneity concern.

Infertility is a valid instrument if it is unrelated to omitted variables that influence labor force participation. Table 1 presents evidence on the validity of our instrument. This table reports coefficient estimates for our fertility measures from a series of regressions (indexed by V_i) that in addition to fertility status control for age as follows:

$$V_i = \theta_1 \text{Infertile}_i + \theta_2(1 - \text{Infertile}_i) + \sum_j \rho_j \text{AGE}_{ji} + \eta_i \tag{2}$$

In essence, these regressions ask, controlling for age, whether infertile women are different from their fertile counterparts ($\theta_1 - \theta_2 = 0$). Table 1 shows that for many important outcome variables infertile women mirror their fertile counterparts. They have the same labor force participation rates, have the same childhood background and became sexually active at the same age. Importantly, when asked about their desired fertility (by gender and number) there is no difference by fertility status. Infertility thus mimics an experiment in which nature assigns, to each woman, a random upper bound for the number of children, independent of background and preferences. However, infertile women are more likely to be married, have more education, and have more educated spouses than their fertile counterparts. Thus it will be important to include these variables in our regression analysis. We also find some evidence that health influences fertility: infertile women are more likely to report a miscarriage and they are more likely to be categorized as obese.

[Table 1 about here.]

Additionally, for our instrument to be valid, infertility should be correlated with the number of children a woman has. Table 2 shows the first stage results. Infertility is indeed strongly correlated with the number of children a women has. On average infertile women have one fewer child, 0.5 fewer preschool-aged children, and are 20 percentage points more likely to be childless than their fertile counterparts. The F-tests show that our instrument has sufficient power in all specifications.

[Table 2 about here.]

5 Results

Column (i) of Panel A in Table 3 presents the OLS estimate, which suggests that each additional child decreases labor force participation by 3.3 percent. Column (ii) contains the corresponding IV estimate. It suggests that the effect of children on labor force participation, using the variation in number of children that comes through the infertility channel, is non-existent. This is the main result of the paper. The IV point estimate (0.002) is close to zero and statistically insignificant suggesting that the OLS parameter was overestimated. This is consistent with the case where unobserved variables, such as career ambition, are important factors explaining both female employment and number of children. In column (iii) of Panel A we take advantage of having two separate measures of infertility and use each as an instrument to run a Hansen J-test for overidentifying restrictions. The GMM model confirms our findings and the J-test statistic rejects the null hypothesis that our instruments are valid.

Continuing in Panel A, Model 2 includes additional control variables. If infertility is unrelated to other determinates of labor force participation then their inclusion should not alter our findings. This is confirmed as the results from Model 2 mirror the results from Model 1. In Model 3 we also add information about health (proxied by BMI) to the regression. It is possible that our infertility measure is capturing poor health and that poor health could directly influence labor force participation, thus invalidating our identification strategy. However, the main finding persists in Model 3. Children have a significant and negative impact in the OLS and no impact on labor force participation in either the IV or the GMM specification.

Panel B repeats the above exercise where the number of children under age 6 is the main variable of interest. Perhaps the barriers to labor force participation are higher for women with pre-school aged children than for women with older children. The OLS estimates suggest that this is indeed the case, as each additional pre-school aged child reduces labor force participation by 7.6 percentage points. Once again, however, the IV results suggest that there is no causal relationship between pre-school aged children and labor force participation.

[Table 3 about here.]

Finally, since infertility impacts the fertility of all women, we can investigate the differences between childless women and women with kids. It is possible that the first child has a large impact on work behavior and that subsequent children have a much smaller impact. In Panel C, our main variable asks if a woman has at least one child. OLS estimates suggest that childless women are much more likely to participate in the labor force. However, columns (ii) and (iii) suggest that there is no causal relationship between motherhood and labor force participation. To summarize, the OLS estimates consistently show a significant negative relationship between labor force participation and the presence or number of children, while the IV and GMM estimates show no effect. This result is robust to numerous alternative specifications. We have also considered heterogeneous effects (not shown) by estimating the models separately by education, age, number of children and country and the main results persist in these subsamples.

6 Conclusions

Our paper investigates the relationship between children and labor force participation for women in Latin America. We use a strategy in which nature prevents some women from obtaining their desired fertility levels. Our results suggest that, at least for the population of women who are not actively controlling their fertility, having children is not a barrier for participation in the paid labor force.

These results contrast Cruces and Galiani (2007) who find that women who are induced to have a third child, out of a desire for a balanced sex mix of their children, are less likely to participate in the labor force. Our identification strategy may be applicable to a broader population. In Cruces and Galiani, the local average treatment effect comes from women whose fertility is altered because of their preference for mixed-sex offspring. Our data contains information on women's preferences over the gender composition of their children. It appears as if women who have a preference for mixed-sex composition are systematically different than the population at large. Infertility, on the other hand, impacts women irrespective of their family preferences, which could explain the difference across findings.

Our results provide little evidence in support of the belief that the rise in female labor force participation in Latin America can be attributed to declining family sizes. However, a common factor such as the empowerment of women could be driving both trends. In addition, our findings suggest that policies focusing solely on family planning are unlikely to increase female labor force participation.

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Tables

Table 1: Women's Characteristics by Fertility Status

Women's characteristics (V_i)	Infertile (θ_1)	Fertile (θ_2)	Test: $\theta_1 - \theta_2 = 0$
Works (=1)	.408 (.032)	.397 (.046)	.011 [.031]
Completed more than primary school (=1)	.519 (.077)	.427 (.061)	.092 [2.51]
Never married (=1)	.099 (.019)	.154 (.019)	-.055 [-4.25]
Age at first marriage (years)	18.3 (.405)	17.6 (.245)	.768 [2.24]
Spouse completed more than primary school (=1)	.585 (.084)	.487 (.073)	.097 [2.40]
Ever had a miscarriage (=1)a	.158 (.022)	.125 (.015)	.033 [2.21]
Body mass index (Obese=1)b	.131 (.021)	.061 (.004)	.070 [3.77]
Grew up in urban areas (=1)	.403 (.061)	.357 (.053)	.047 [1.40]
Age at first intercourse (years)	17.47 (.216)	17.2 (.118)	.294 [1.52]
Ideal number of sons	.975 (.100)	1.00 (.064)	-.028 [-0.41]
Ideal number of daughters	1.06 (.139)	1.03 (.087)	0.025 [0.31]

Notes: a Information on miscarriages not available for Bolivia in 1998. b Information on BMI not available for the full sample. Standard errors in parenthesis and t-statistics in brackets.

Table 2: First Stage Results of Infertility on Number of Children

Dep. Var.:	Children at Home			Children Under 6			Has at Least One Child		
Model:	1	2	3	1	2	3	1	2	3
Coefficient	-1.093	-1.109	-0.976	-0.507	-0.532	-0.408	-0.212	-0.227	-0.160
Std. Error	(0.064)	(0.069)	(0.068)	(0.039)	(0.041)	(0.035)	(0.012)	(0.010)	(0.029)
F-test	292.3	260.5	207.9	170.9	168.4	138.6	338.8	499.9	29.8
Obs.	24,131	24,131	24,131	24,131	24,131	24,131	15,992	15,992	15,992

Note: Robust standard errors clustered at regions in parenthesis. Model 1 includes: age categories, education categories, age and education interactions and country fixed effects. Model 2 includes Model 1 and marital status, age at first marriage, spouse's educational categories, childhood place of residence, and age at first intercourse. Model 3 includes Model 2 and BMI indicators. All models include sample weights.

Table 3: The effect of children on women's labor force participation

	Dependent variable: Women's labor force participation (LFP_i)								
		Model 1			Model 2		Model 3		
	OLS	IV	GMM	OLS	IV	GMM	OLS	IV	GMM
	(i)	(ii)	(iii)	(iv)	(v)	(vi)	(vii)	(viii)	(ix)
Panel A. K_i: Number of children at home									
β	-0.033	0.002	0.008	-0.026	-0.005	-0.002	-0.024	-0.003	-0.006
s.e.(β)	(0.003)	(0.018)	(0.015)	(0.003)	(0.015)	(0.013)	(0.004)	(0.018)	(0.018)
J-statistic			[0.70]			[0.90]			[0.09]
Panel B. K_i: Number of children under 6									
β	-0.076	0.005	0.016	-0.061	-0.011	-0.005	-0.057	-0.006	-0.005
s.e.(β)	(0.006)	(0.038)	(0.036)	(0.005)	(0.032)	(0.030)	(0.007)	(0.043)	(0.046)
J-statistic			[0.65]			[0.92]			[0.09]
Panel C. K_i: Has at least one child									
β	-0.1	0.012	0.037	-0.067	-0.026	-0.01	-0.099	-0.016	-0.023
s.e.(β)	(0.012)	(0.091)	(0.071)	(0.009)	(0.076)	(0.061)	(0.013)	(0.111)	(0.108)
J-statistic			[0.71]			[0.90]			[0.09]
Obs.	24,131	24,131	24,131	24,131	24,131	24,131	15,992	15,992	15,992

Notes: Robust standard errors clustered at regions in parenthesis. P-value for Hansen J statistic in brackets. See Table 2 for model definitions. All models include sample weights.