

Education Gains Attributable to Fertility Decline: Patterns by Gender, Period, and Country in Latin America and Asia

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Abstract We investigate the heterogeneity across countries and time in the relationship between mother’s fertility and children’s educational attainment—the quantity-quality (Q-Q) trade-off—by using census data from 17 countries in Asia and Latin America, with data from each country spanning multiple census years. For each country-year, we estimate micro-level instrumental variables models predicting secondary school attainment using number of siblings of the child, instrumented by the sex composition of the first two births in the family. We then analyze correlates of Q-Q trade-off patterns across countries. On average, one additional sibling in the family reduces the probability of secondary education by 6 percentage points for girls and 4 percentage points for boys. This Q-Q trade-off is significantly associated with the level of son preference, slightly decreasing over time and with fertility, but it does not significantly differ by educational level of the country.

Keywords Quantity-quality trade-off · Fertility and education · Cross-country comparison · Instrumental variables · Heterogeneous effects

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Introduction

Has the worldwide decline in birth rates improved the lives of newer generations? Education is a catalyst in this relationship. Reduced cohort sizes could facilitate governments in improving human capital by increasing the coverage and quality of public education. At the household level, reduced family size could allow parents to allocate more resources to the education of each child. In combination, these responses to lower birth rates have the potential to drive higher economic growth (Lutz et al. 2008). The goals of our study are to test the hypothesis that reduced family size improves education of children in the broadest possible variety of settings and times and to determine whether this effect differs by gender, period, and country.

A substantial empirical literature has examined theories of quantity-quality (Q-Q) trade-offs that parents and governments make between the number of children and the amount of investment in the quality of those children—for example, as posited by Becker and Lewis (1973). Earlier evidence consistent with such trade-offs had been found in a variety of settings (e.g., Blake 1981 and Hanushek 1992 in the United States; Knodel and Wongsith 1991 in Thailand; and Rosenzweig and Wolpin 1980 in India), suggesting that fertility decline can be an important precursor of increasing investments in children's human capital, such as their educational attainment. More recent studies in other contexts have confirmed such relationships in developing countries such as China (Li et al. 2008; Rosenzweig and Zhang 2009), South Korea (Lee 2008), and Costa Rica (Li et al. 2014). However, several other recent studies in wealthier countries have called into question the universality of such relationships: Black et al. (2005) in Norway, Åslund and Grönqvist (2010) in Sweden, and Angrist et al. (2010) in Israel all found null effects of fertility on schooling.

Comparing these results across prior studies is challenging because of variation in methodologies, periods, and country settings. In this article, we apply a common methodology to compare the estimated effects of number of siblings on secondary school attainment across many developing countries and periods. The method that we apply is instrumental variables (IV), using sex composition of the first two children as identifying instruments. As with all methods, there are potential threats to the validity of this approach (we also present ordinary least squares (OLS) results to better understand robustness), but it has been applied in numerous studies previously (e.g., Jensen 2005; Lee 2008; Li et al. 2014) and allows us to standardize one aspect of the studies while examining how results vary by other important aspects of the settings. We discuss in more depth some of the concerns with sex composition as instruments in the [Discussion](#) section.

The first pattern that we test for is whether there is any systematic trend toward smaller (or larger) Q-Q trade-offs over time. Average estimated effects in the research literature do appear to be decreasing over time, but it is unclear whether this is due to the effects of methods, different countries being tested, or true changes in underlying behavioral trade-offs. A variety of factors could be hypothesized to cause changing effects over time, but because of unclear and potentially competing effects, we do not have an *a priori* hypothesis regarding how these might shape overall time trends. To better understand such possibly changing relationships, we next examine in turn the patterns of results by four time-varying country-level characteristics: changing gender preferences, decreased fertility, increased wealth, and increasing average educational

attainment. Consistent with prior literature, we also examine all relationships separately for boys versus girls.

Finally, to help interpret the levels and variation in magnitudes of our estimated Q-Q trade-offs, we present simulations of how much of the increase in secondary school enrollment over time could be statistically explained by decreased fertility in each setting if our estimates were interpreted causally.

We focus on countries from two regions—Asia and Latin America—for several reasons. First, both regions mostly comprise developing countries to which we limit our sample such that the countries we examine are more comparable. Literature also shows that Q-Q trade-off is likely to be more prominent in developing countries (Li et al. 2014). Second, Asian countries are generally considered to have strong son preferences, while the Latin American region is thought to have generally balanced gender preferences (Arnold 1997). This distinction is important because we use gender composition to instrument for number of siblings and because we use gender preference to directly predict the estimated Q-Q trade-off in each country-year. Thus, having countries with varying gender preferences in our sample is meaningful both theoretically and methodologically. Third, we exclude African countries in the study because the complexity in their household living arrangements precludes the use of simple census data to establish the number of siblings of each child, as detailed in the [Methods and Data](#) section.

It is crucial to note that we do not attempt to representatively characterize Q-Q trade-off in either Asia or Latin America. Instead, the goal is to gain further insight in Q-Q trade-off through comparison across a set of countries with reasonable similarity and diversity. In particular, we attempt to explore any systematic time trend in Q-Q trade-off, which requires at least two years of data for each country. We provide more details on the selection criteria of countries in our sample in the next section.

Methods and Data

Analysis Strategy

We first estimate the relationship between number of siblings and educational attainment in as many Latin American and Asian developing countries and years as possible with micro-level census data, estimating separate models for each census country-year. In a second stage of country-year-level analysis, we examine the variation across countries and years in the siblings-education association.

Individual-Level (Micro-Level) Analysis

Following the methodology presented in Li et al. (2014), we estimate two-stage least squares (2SLS) versions of IV models separately for each country-year in micro-data census samples. These models predict the probability of having any secondary school education among children aged 15–17 as a function of the number of siblings they have.

To overcome the likely endogeneity of fertility, we use the sex composition of the mother's first two births as instruments for the child's number of siblings in the IV

model. Angrist and Evans (1998) first used this approach to analyze parents' labor supply. Since then, sex composition indicators have been used as IV in numerous studies (Jensen 2005; Lee 2008; Qian 2004) to examine the effect of sibling size on children's education. The first-stage explanatory power of these instruments depends on the extent to which sex composition of the elder children alters fertility decisions because of parental preferences for having boys versus girls. Because preference on the sex of offspring is likely to be context-specific, we do not attempt detailed discussion of such preferences in all the countries we examine. However, as discussed earlier, previous literature has established that many Asian countries have preferences for having sons (Das Gupta et al. 2003; Edlund 1999; Lee 2008; Qian 2004), whereas countries in Latin America have a general preference of having a balanced sex composition among children, with a slight preference for having girls (Bongaarts 2001; Li et al. 2014). Despite the potential heterogeneity in parental response to the sex composition of the elder children, it is key for comparability purposes in this study to adopt the same methodology across countries and years. In the upcoming Discussion section, we elaborate on the potential implications for interpreting our results across heterogeneous settings and discuss other potential weaknesses of our identification strategy.

An important decision involved in the IV strategy is the number of children on which sex composition should be computed. We compute the sex composition only for the first two births because the choice between having two or three children is a relevant decision for a substantial fraction of families in all the countries in our sample. (We present the summary statistics on sibling size in the Results section.) In addition, one may be more concerned about endogeneity in a three-birth instrument if selection on sex composition occurs between the second and the third birth, thereby justifying our choice of using sex composition of the first two births. We operationalize sex composition as a vector of two binary variables, with the first indicating whether the first two births in the family are both boys, and the second indicating whether the first two births are both girls. Accordingly, our final analytical samples are restricted to those children with at least one sibling in the household.

Because sex composition of the first two births is mostly likely to affect a couple's decision of having a third birth, we estimate the 2SLS models only on first and second births who are the relevant group of children affected by the instrument: that is, children who are the third births or above potentially may not have been born had the sex composition of the first two births been different. We estimate the models separately for boys and girls in addition to the full sample to capture any systematic differences in the estimated relationship by gender.

Model identification relies crucially on the assumption that sex composition affects educational attainment only through family size (number of siblings). To minimize concerns arising from possibly differential investment for girls who have only brothers (as discussed by Butcher and Case 1994), the models also control for whether the child has any sisters.

For comparison, we also estimate the same models via OLS for each country-year—that is, by regressing a child's secondary education attendance on the observed number of siblings. These models are reported in the Online Resource 1.

Country-Year–Level (Macro-Level) Analysis

A country-year-level analysis is necessary to make sense of the many estimates we obtain from the micro-level models. In essence, we are interested in explaining the variation in the effect of number of siblings on education (as estimated by IV coefficients) in the micro-level models using country-year-specific characteristics. This analysis helps us understand the contextual factors that may shape the effect of having an additional sibling on children's educational attainment.

Because our dependent variable in the macro-level model is the estimated country-year-specific IV coefficient, simple OLS is no longer efficient because the dependent variable is itself observed with varying sampling errors, which then lead to heteroskedastic errors in the macro model (Hanushek 1974; Saxonhouse 1976). To address this issue, we follow Saxonhouse (1976) in weighting each country-year observation in the macro-level model by the inverse of the estimated standard error of the dependent variable, which is the estimated standard error of the IV coefficient in the micro-level model.

Given the panel structure of the country-year-level data, a natural choice is a weighted fixed-effects (FE) model with country fixed effects. This model relaxes the assumption that the country-specific effects are uncorrelated with the country-year-specific characteristics included as independent variables. On the other hand, the FE model uses only within-country variation to estimate coefficients of interest and can be less efficient than a weighted least squares (WLS) model if the preceding assumption plausibly holds.

To determine the appropriate choice of model, we estimate both a WLS model and a weighted FE model with country fixed effects and conduct a cluster-robust version of the Hausman test (Cameron and Miller 2015) for the difference between the estimated coefficients. If the test suggests that the WLS and weighted FE estimates are not significantly different, the WLS model is superior given that it is more efficient.

Data

Data used in this study come from 2 % to 10 % of the censuses samples publicly available from the University of Minnesota Population Center (2014) IPUMS-International web dissemination system. After creating the analytical sample for each country-year (which we discuss in more detail shortly), we keep only those countries and years with at least 5,000 child observations, meeting our inclusion criteria to allow for meaningful statistical inference. As mentioned earlier, we also require that each country have at least two years of available data in our final sample in order to examine any systematic time trends. After implementing these restrictions, our final analytical sample consists of 5 countries in Asia^{1,2} and 12 countries in Latin America, with a total of 48 country-years covering the period 1970–2010. Table 1 shows the complete list of country-years as well as the number of observations in our final analytical sample.

¹ IPUMS generally has fewer number of years data available for each country in Asia compared with Latin America, which limits the number of Asian countries that we include in our data. We also exclude China from our sample because of its one-child policy, which artificially restricted the number of children in each family.

² We include Turkey as part of Asia for the purpose of this study despite the country being a member of the European Union.

Table 1 Country-year characteristics and values of estimated coefficients on number of siblings in micro-level IV model (IV $\hat{\beta}$) predicting the probability of any secondary school for all country-years

Country	Year	<i>N</i>	% Any Secondary School		Number of Siblings: All	% Last Birth Male: All	IV $\hat{\beta}$		
			Boys	Girls			All	Boys	Girls
Cambodia	1998	19,853	0.360	0.276	3.780	0.508	-0.027 [†]	-0.028	-0.019
Cambodia	2008	27,315	0.572	0.588	2.728	0.507	-0.040***	-0.034*	-0.045**
Indonesia	1980	65,937	0.435	0.377	3.610	0.511	-0.016**	-0.005	-0.032***
Indonesia	1990	12,173	0.647	0.634	3.118	0.521	-0.046**	-0.053**	-0.045 [†]
Indonesia	2000	358,807	0.516	0.558	2.284	0.501	-0.030***	-0.024***	-0.041***
Thailand	1970	9,239	0.391	0.352	4.229	0.525	-0.057***	-0.048*	-0.071 [†]
Thailand	1980	5,984	0.295	0.271	3.445	0.514	-0.097***	-0.109**	-0.056
Thailand	1990	9,455	0.513	0.524	2.073	0.516	-0.056**	-0.043 [†]	-0.077**
Thailand	2000	14,022	0.836	0.902	1.479	0.518	-0.061***	-0.065**	-0.057*
Turkey	1985	24,085	0.464	0.400	2.888	0.544	-0.116***	-0.097***	-0.198***
Turkey	1990	40,570	0.495	0.424	2.745	0.551	-0.127***	-0.115***	-0.255***
Turkey	2000	59,780	0.572	0.501	2.247	0.551	-0.082***	-0.073***	-0.151***
Vietnam	1989	38,821	0.765	0.751	3.359	0.543	-0.056***	-0.048**	-0.097**
Vietnam	1999	59,016	0.833	0.814	2.306	0.558	-0.100***	-0.095***	-0.147***
Vietnam	2009	399,726	0.926	0.931	1.655	0.569	-0.046***	-0.044***	-0.083***
Argentina	1970	6,735	0.580	0.625	2.065	0.523	-0.084***	-0.075***	-0.108***
Argentina	1980	35,822	0.607	0.679	2.384	0.512	-0.067***	-0.068***	-0.066***
Argentina	2001	73,470	0.882	0.923	2.290	0.510	-0.031***	-0.033***	-0.026***
Bolivia	1992	6,880	0.838	0.789	3.229	0.506	-0.044*	-0.039 [†]	-0.041
Bolivia	2001	6,868	0.772	0.788	2.755	0.493	-0.038**	-0.023	-0.062**
Brazil	1980	79,403	0.137	0.181	3.304	0.510	-0.016***	-0.012*	-0.021**
Brazil	1991	133,457	0.130	0.194	2.655	0.508	-0.023***	-0.015***	-0.035***
Brazil	2000	200,022	0.266	0.371	2.038	0.509	-0.034***	-0.024***	-0.049***
Brazil	2010	181,808	0.378	0.500	1.756	0.511	-0.044***	-0.036***	-0.056***
Chile	1982	15,889	0.618	0.670	2.513	0.504	-0.015	-0.008	-0.033
Chile	1992	23,197	0.764	0.804	1.977	0.510	-0.034**	-0.032*	-0.040*
Chile	2002	29,363	0.850	0.889	1.754	0.507	-0.024**	-0.025*	-0.022*
Colombia	1973	15,921	0.522	0.558	4.367	0.511	-0.023*	-0.026*	-0.023
Colombia	1985	30,013	0.645	0.711	2.926	0.505	-0.023**	-0.016	-0.030**
Colombia	1993	48,359	0.741	0.802	2.461	0.509	-0.041***	-0.038***	-0.040***
Dominican Republic	2002	11,060	0.612	0.749	2.191	0.509	-0.079***	-0.066**	-0.099**
Dominican Republic	2010	14,857	0.704	0.834	2.025	0.502	-0.013	0.000	-0.039
Ecuador	1982	6,041	0.647	0.695	3.674	0.503	-0.050*	-0.039	-0.115*
Ecuador	1990	11,385	0.764	0.786	3.004	0.506	-0.011	-0.013	-0.013
Ecuador	2001	16,188	0.713	0.745	2.396	0.509	-0.030*	-0.053***	0.007
Ecuador	2010	24,238	0.849	0.882	2.138	0.506	-0.036***	-0.027*	-0.052***

Table 1 (continued)

Country	Year	<i>N</i>	% Any Secondary School		Number of Siblings: All	% Last Birth Male: IV $\hat{\beta}$		Boys	Girls
			Boys	Girls		All	All		
El Salvador	1992	5,792	0.237	0.328	2.468	0.515	-0.030 [†]	-0.011	-0.043 [†]
El Salvador	2007	10,672	0.231	0.275	2.341	0.507	-0.004	-0.002	-0.008
Mexico	1990	129,727	0.777	0.791	3.166	0.510	-0.050***	-0.045***	-0.060***
Mexico	2000	164,745	0.816	0.831	2.680	0.509	-0.041***	-0.032***	-0.059***
Mexico	2010	234,679	0.868	0.893	2.358	0.513	-0.029***	-0.022***	-0.041***
Nicaragua	1995	5,171	0.474	0.599	3.523	0.510	-0.042*	-0.068**	-0.011
Nicaragua	2005	7,219	0.633	0.752	2.579	0.511	-0.055**	-0.049 [†]	-0.078**
Peru	1993	25,007	0.869	0.867	2.812	0.503	-0.031***	-0.023**	-0.042*
Peru	2007	42,746	0.910	0.914	2.157	0.509	-0.026***	-0.028**	-0.024**
Venezuela	1971	12,125	0.431	0.474	3.933	0.510	-0.031**	-0.040**	-0.018
Venezuela	1990	21,685	0.620	0.734	3.015	0.500	-0.035***	-0.043**	-0.021
Venezuela	2001	36,001	0.815	0.895	2.024	0.506	-0.030***	-0.030**	-0.032**

Notes: Coefficients shown in columns 3–5 are the estimated two-stage least squares (2SLS) coefficients on number of siblings predicting the probability of any secondary school, using the sex composition of the first two births in the family as instruments on the sample of girls only. See an example of the 2SLS models run for Cambodia in Tables S3 (first stage) and S4 (second stage) in Online Resource 1. The 2SLS models for all country-years have first-stage partial *F* statistics on the instruments greater than 10.

[†]*p* < .10; **p* < .05; ***p* < .01; ****p* < .001

Establishing the number of siblings of a child in a census database is not trivial because there are no direct survey questions about it, and neither does census collect birth-history data, which would allow direct calculation of number of children that each mother has. The IPUMS data contain a variable that identifies the mother of a given individual if the mother also lives in the same household as the individual. This variable does not come from the census questionnaire but rather is constructed using a set of algorithms and assumptions. Although not perfect, this is the best information we have in the IPUMS data on mother-child link, which allows us to estimate the number of siblings of an individual as long as the mother lives in the same household. The limitation to this approach, besides the potentially imperfect mother-child link, is that we could examine only those families in which all children live with their mother.

Because children are more likely to move away from their parents after they become adults, we restrict our sample to children aged 15 to 17 to limit the possible selection effect in our final analytical sample. The lower bound on the age limit guarantees that the children are eligible for at least the first year of secondary school: children usually start secondary school at approximately age 13 or 14. Obviously, the age restriction does not completely remove the potential bias caused by the requirement that all children live with their mother in the same household. For instance, such households may be more financially stable and affluent than those in which the children move

away from their mother at a young age. Further analysis suggests that more than 80 % of children aged 15 to 17 in the census data live with their mothers in the vast majority of countries we examine,³ which somewhat alleviates the concern for selection bias.

We derive our analytical sample for each country and year as follows. First, we identify mothers of the children in our final sample by imposing the age range of 25 to 70 among females in the census data. This restriction prevents including mothers who have children when they are either too young or too old and minimizes age-related measurement errors. Second, we identify all children in the data who are linked with the mothers, retaining only children of mothers who have at least two children and whose children are all reported in the household (i.e., whose constructed number of children matches the count of children in the data). Third, we estimate the mother's age at first birth by subtracting from her age the age of her oldest child; we exclude mothers (and their children) who were calculated to be younger than age 10 at their first birth. Our final analytical samples consist of children aged 15 to 17 who are the first or second birth of the mothers meeting all the preceding restrictions. Table 1 lists the number of observations (children) in our final analytical sample for each country-year.

Measures

Our measure of educational attainment, which is our outcome of interest, is whether the child has attended at least one year of secondary school. We have two primary reasons for using this measure of educational attainment instead of other alternatives. First, continuing education after completing elementary school is an important decision made by families in the majority of countries that we examine. (We show the country-year-level summary statistics on percentage of children with any secondary education in the [Results](#) section.) Second, the restriction of children living with their mothers and the corresponding upper age limit makes it necessary to use education measures experienced sufficiently early in the schooling system.

The key explanatory variable of interest is the number of siblings the child has. We obtain this measure by subtracting 1 from the mother's constructed number of children alive. As mentioned earlier, we keep a child in the sample only if her mother's constructed number of children matches the count of children who are present in the household and linked to the same mother. As instruments for number of siblings, we use the sex composition of the first two births in the same family.

As potential confounders of the association between number of siblings and education, we include in the micro-level regressions the following covariates: child's gender (a dummy variable for female), child's age dummy variable, child's birth order (a dummy variable for first birth, with second birth as the omitted category), whether the child has any sisters in the family, mother's schooling (a dummy variable for each level of education being the highest level attained), mother's age at first birth, and an index of family asset.

In addition, we define a *region* variable in each country as the highest level of geographical area available in the census data that guarantees at least 30 regions. For example, in a census with 10 provinces and 40 municipalities, our region variable will be the municipality. We cluster the standard errors in the micro-level models using the

³ Li et al. (2014) found that at least in the Costa Rican context, results were insensitive to alternative age choices.

region variable. We also include in the regressions a series of dummy variables indicating the region in which the family is located. Accordingly, we use only within-region and within-cohort variation to identify the effect of fertility on education.

In the country-level analysis, we use several explanatory variables in predicting the relationship between siblings and education. These variables include the country-year means of the outcome variable and the key explanatory variable in the micro models: namely, the proportion of children with secondary schooling and the mean number of siblings. We also include as explanatory variables the extent of son preference (measured as the percentage of youngest children who are male) and a measure of mean household wealth for each country-year created as follows. First, we assign 1 to each amenity that the household owns or 0 for each that it does not own from a list of amenities commonly used to create the asset index: running water, toilet, refrigerator, TV, telephone, washing machine, and car. Second, we tally the total number of amenities owned by each household and divide it by the maximum number of amenities owned by households in the same country and year. Third, we take the average of the household-level asset index across all households in the same country and year, which is our final measure for wealth at the country-year level.

Results

Summary Statistics and Micro-Level Results

Table 1 presents the country-year-level data: the means for the key dependent and explanatory variable in each micro-level model, which are also used as covariates in the macro-level analysis. Table 1 also contains the estimated IV coefficients in each micro-level regression indicating the effect of number of siblings on the probability of having any secondary education. The IV coefficient estimate can be interpreted as the change in the probability of having at least one year of secondary schooling in response to having one more sibling, for children who are either the first or second birth in the family. The IV coefficients are negative in all but 1 of 96 regression models that were separately estimated for boys and girls, implying that schooling diminishes with additional siblings. Moreover, the majority of the coefficients are statistically significant, with small standard errors. Table S1 of Online Resource 1 presents the analogous coefficient estimates from OLS models of the effect of number of siblings on the probability of any secondary school, including the same covariates as are included in the second stage of the IV model. The OLS coefficients also have a negative sign in all but 2 of the 96 regression models.

Table S2 (Online Resource 1) presents the first stage coefficients on the instruments, which are the dummy variables indicating whether the first two births are both boys (girls), for regressions on the full sample and by gender. It also presents the partial F statistics on the instruments; the F statistics are uniformly large, indicating that the instruments have sufficient explanatory power in the first stage. For illustrative purpose, Tables S3 and S4 in the online supplement present the complete first-stage and second-stage micro-level models and coefficient estimates for one country: Cambodia. Although the first-stage F statistic is important for identification purposes, the signs of the first-stage coefficients on the instruments are difficult to interpret because these coefficients are conditioning on whether the child has any sisters in the family, which is included as

another covariate to control for differential parental investment, as discussed by Butcher and Case (1994). To more easily interpret the fertility effects of “first two boys” and “first two girls” on fertility, we also include in Table S2 (final three columns) the results of analogous country-specific models regressing household-level fertility as a function of sex composition of the first two children. These results reveal that sex composition is related to fertility but in different ways in different countries. Some country-years show strong son preference (the “gg” coefficient for the first two births being girls is strongly positive and significantly larger than the analogous “bb” coefficient for the first two births being boys). Those country-years for which the p values in the final column reveal larger effects for girls are mostly in Asia, which is consistent with the general impression of son preference there. In many country-years, though, there appears to be a strong preference for mixed genders: fertility is similarly higher when the first two children are of the same gender regardless of whether they are boys or girls. Preference for mixed gender is apparent in many Latin American countries but also in several Asian countries. For example, Cambodia, Indonesia, and Thailand do not consistently exhibit strong son preference in our data, consistent with previous literature (Kevane and Levine 2000; Wongboonsin and Ruffolo 1995; Zimmer and Kim 2001).

Table 2 presents the means and standard deviations (across country-years) of the micro-level OLS and IV coefficients. The (unweighted) mean IV coefficient across country-years is -0.044 , which is quite close to the mean coefficient estimated on boys

Table 2 Summary statistics of country-year-level variables, by sample

	All (1)	Boys (2)	Girls (3)
Estimated Coefficient on Number of Siblings in Micro-Level Models			
OLS	-0.033 (0.015)	-0.032 (0.016)	-0.033 (0.016)
IV	-0.044 (0.027)	-0.041 (0.027)	-0.058 (0.049)
Explanatory Variables			
Percentage last birth male	0.515 (0.016)	—	—
Mean number of siblings	2.686 (0.681)	—	—
Percentage any secondary school	0.626 (0.216)	0.611 (0.215)	0.642 (0.219)
Proportion of wealth	0.528 (0.317)	—	—
Number of Countries	17	17	17
Number of Country-Years	48	48	48

Notes: Main coefficients reported are means across country-years. Standard deviations of the means are shown in parentheses. For percentage last birth male, mean number of siblings, and wealth proportion, we use the overall sample value as regressors in the by-gender analysis and thus do not report the by-gender subsample values here.

only (-0.041) and slightly smaller in magnitude than the mean IV coefficient for girls only (-0.058). The IV coefficient estimates are, on average, approximately one-third larger than the OLS estimates, albeit with more variation. In micro-models estimated on girls only, the IV coefficients are nearly twice as large as the OLS coefficients.

Table 2 also reports the summary statistics of the four country-year-level covariates used in the macro-level regressions to explain the variation in the estimated micro-level coefficients. For proportion with any secondary school, we report the summary statistics over the full sample as well as by gender given that meaningful difference can exist in secondary school attainment between boys and girls. Accordingly, in the macro-level models, we control for gender-specific secondary school attainment when estimating models separately for boys and girls. (The other covariates are measured based on the full sample rather than gender-specific means.)

The mean proportion of male last births across all country-years in our sample is 0.515, representing a relatively balanced male-to-female ratio among last births. This finding is perhaps not surprising because most of the countries we examine are Latin American, with generally balanced gender preference. The mean of the country-year-specific means of number of siblings of children in our sample is approximately 2.7, indicating that families have an average of 3.7 children (conditional on our sample inclusion criteria of having at least two children). Across country-year samples, the mean percentage of children with any secondary schooling is 62.6 %, with girls (64.2 %) higher than boys (61.1 %) on this measure. In 13 countries, this measure is below 70 % in one or more years in our data, confirming that continuing education after completing elementary school is likely an important margin for decision by families. Finally, the typical family in our sample owns approximately one-half of the amenities used to calculate the wealth measure, albeit with substantial variation (in both the means and in the number and types of assets measured).

Graphical Analysis of Macro-Level Patterns

To better understand the relative magnitude of the coefficients and covariates, we first display macro relationships graphically. Figure 1 plots the IV coefficient estimates averaged across years for each country, separately by gender. For all countries except Venezuela and Nicaragua, the mean IV coefficient for girls is more negative than that for boys, indicating that girls' education is in general more susceptible to the influence of having an additional sibling than is boys' education. Plotted mean coefficients for both boys and girls vary more among Asian countries than among Latin American countries. This variation is driven primarily by two countries, Turkey and Vietnam; thus, we report later herein sensitivity to dropping these countries from the macro-level regressions.

Figure 2 adds the time dimension and graphs country-year-specific IV coefficients on the y -axis against year on the x -axis. We do this separately for coefficients estimated on the full sample (i.e., combining genders) and then separately by gender. We also plot linear regression lines that represent time trends for each region and sample. In general, the IV coefficients become smaller (less negative) across time, although the slope of the fitted line is quite small in all cases. The fitted lines for Asian and Latin American countries are almost parallel, suggesting similar time trends. However, Asian countries

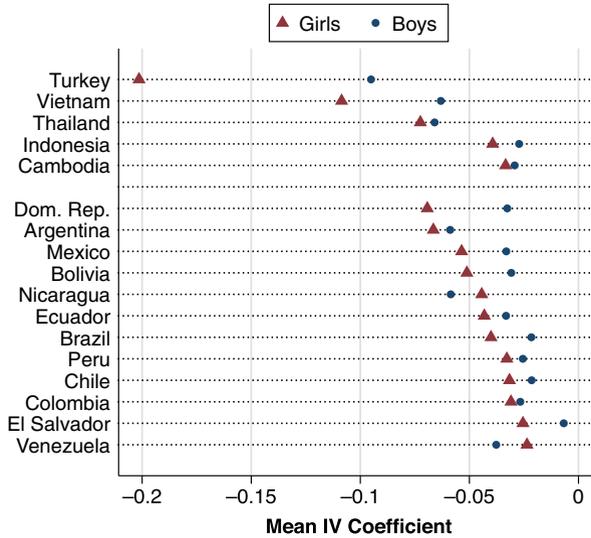


Fig. 1 Mean estimated coefficients on number of siblings in micro-level IV models by country and sample. Coefficients plotted are the estimated 2SLS coefficients at the country-year level on number of siblings predicting the probability of any secondary school, using the sex composition of the first two births in the family as instruments, then averaged across years within country. See Table 1 for the complete list of country-years and the corresponding 2SLS coefficients

have more negative intercepts (larger relationships) than those for Latin American countries, suggesting persistent level difference in the relationship between number of siblings and secondary school attendance (which is the largest for girls).

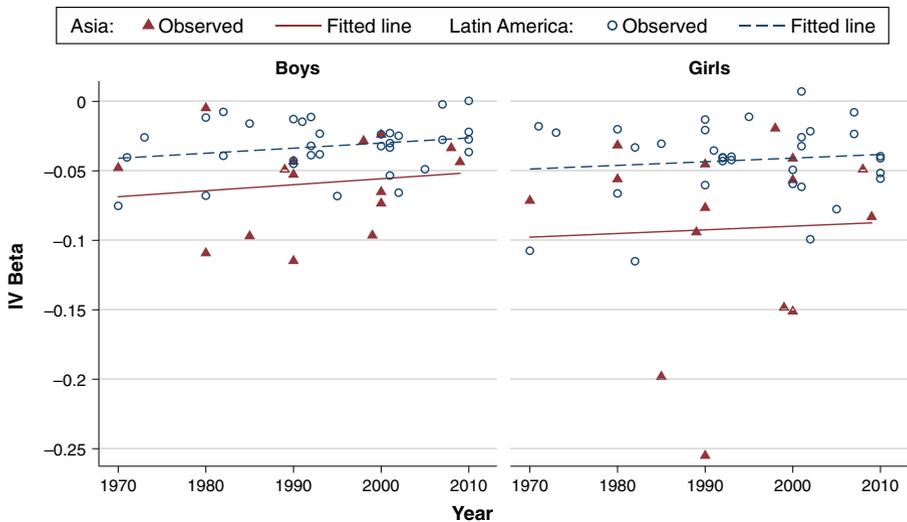


Fig. 2 Scatterplots and fitted lines of time trends in estimated coefficient on number of siblings in micro-level IV models, by sample. Values plotted are the estimated 2SLS coefficients at the country-year level on number of siblings predicting the probability of any secondary school, using the sex composition of the first two births in the family as instruments, by gender. The slope of the fitted lines indicates time trend in the estimated 2SLS coefficients for Asia and Latin America, respectively

In Fig. 3, we present similar trend graphs for the four country-year-level explanatory variables to understand how these covariates are changing over time. Panel a plots percentage last birth male as a function of time. Interestingly, there seems to be an upward trend in percentage last birth male in Asian countries; this upward trend, coupled with the downward trend in mean number of siblings plotted in panel b, could mean that more families in Asia follow the son-stopping rule (stopping having children after a son is born) when general fertility and hence the chance of having at least one son is lower. Latin American countries, on the other hand, have remained relatively stable in percentage last birth male over time, at levels lower than in Asian countries, and with a slight negative trend. Perhaps unsurprisingly, both mean asset index and percentage children with any secondary schooling exhibit conspicuous upward trends among both Asian and Latin American countries, as indicated by panels c and d of Fig. 3.

Finally, Fig. 4 presents a series of graphs with IV coefficients plotted on the y -axis and country-year explanatory variables plotted on the x -axis. Panel a has percentage last birth male as the x -axis variable. Regardless of sample, IV coefficients become more negative as percentage last birth male increases, and this is the case for both Asian and Latin American countries. By contrast, the IV coefficients do not seem to exhibit similarly prominent or consistent relationships with the other three variables (mean number of siblings, asset index, and percentage any secondary school): the fitted lines are all relatively flat. Exceptions are that the IV coefficient seems to become slightly

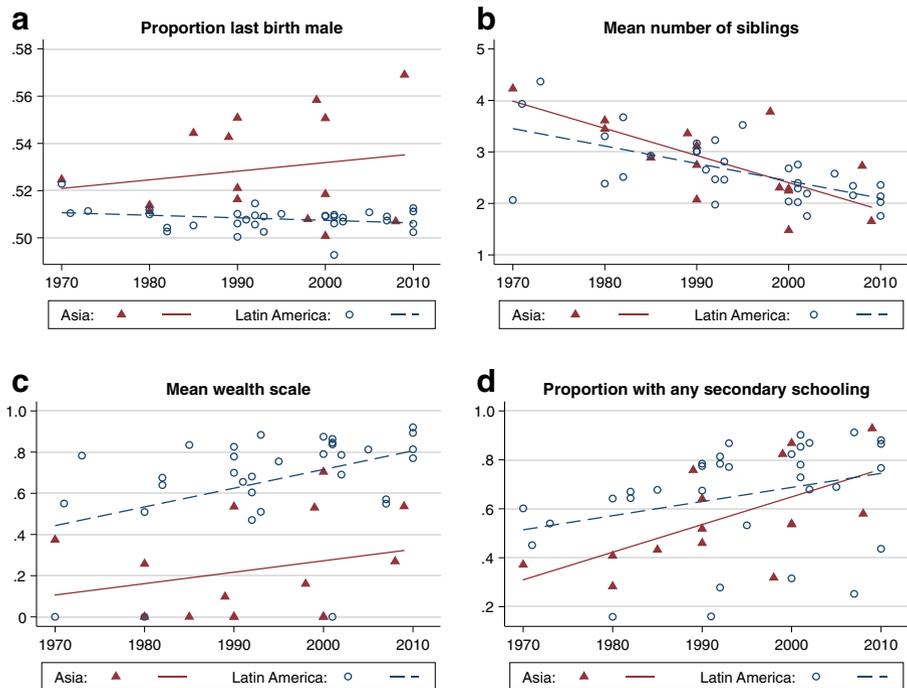


Fig. 3 Scatterplots and fitted lines of time trends in country-year characteristics. Values plotted are means of the country-year-specific characteristics. The slope of the fitted lines indicates time trend in the plotted characteristics for Asia and Latin America, respectively

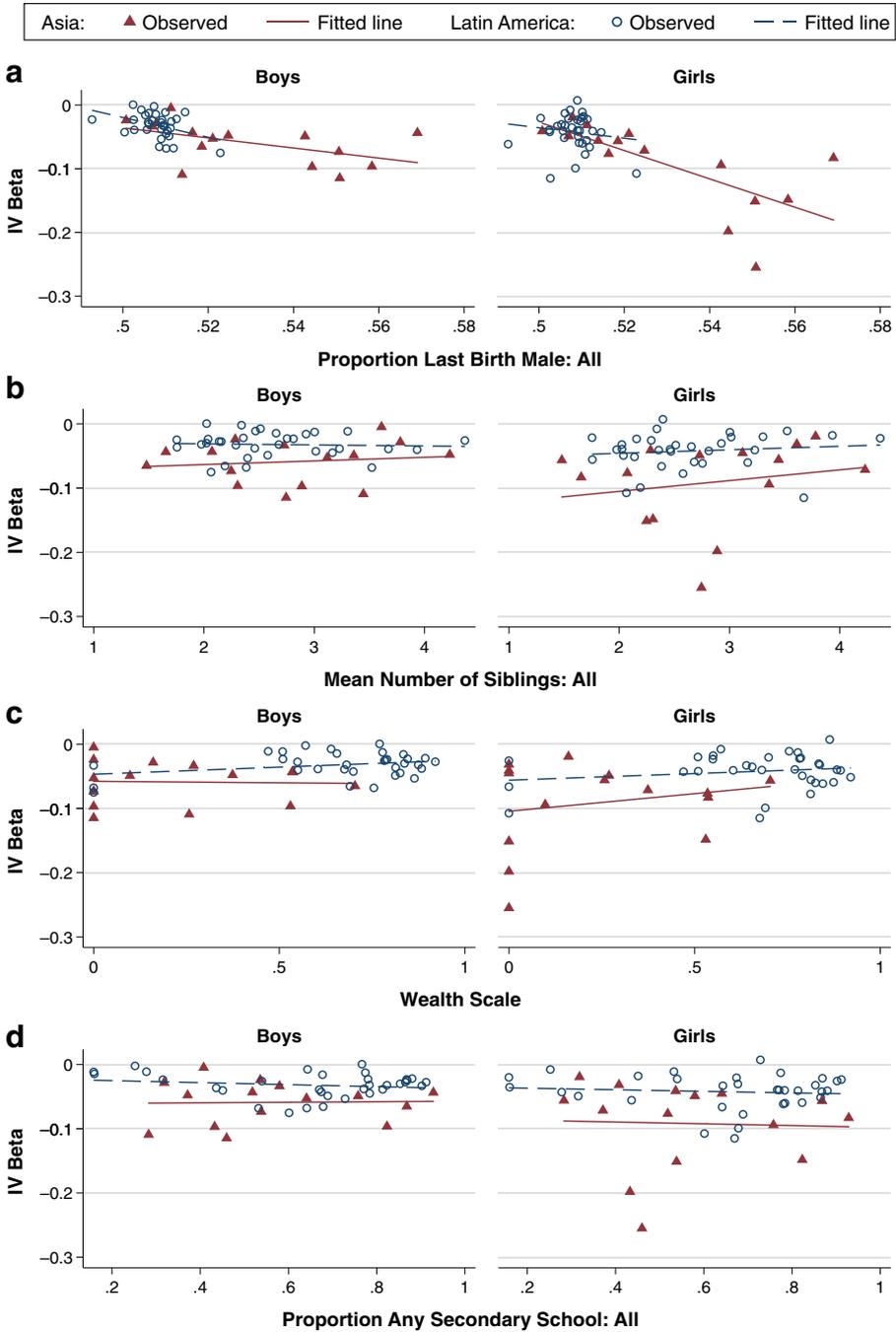


Fig. 4 Scatterplots and fitted lines of IV on various country-year characteristics, by sample. Values plotted are the estimated 2SLS coefficients at the country-year level on number of siblings predicting the probability of any secondary school, using the sex composition of the first two births in the family as instruments, by gender. The 2SLS coefficients are plotted against the means of the country-year-specific characteristics. The slope of the fitted lines indicates the predicted relationship for Asia and Latin America, respectively

less negative when the mean number of sibling increases in Asian countries (panel b) and as the asset index increases in Latin American countries (panel c).

For comparison, we present equivalent graphs in Online Resource 1 for the OLS coefficients. The general pattern exhibited in these graphs is very similar to that for IV coefficients, although the supplemental graphs show much less variation in OLS coefficients across countries and between genders.

Macro-Level Regression Results

Finally, in Table 3, we present results from the macro-level regression models that predict the micro-level IV coefficients using the aforementioned country-year characteristics. We again present the results for the full sample and separately by gender.

Table 3 Country-year-level weighted least squares (WLS) regressions of estimated coefficients on number of siblings in micro-level IV model predicting Q-Q trade-offs using country-year characteristics, by sample

	All (1)	Boys (2)	Girls (3)
Percentage Last Birth Male	-0.855* (0.355)	-0.739* (0.322)	-1.809** (0.597)
Mean Number of Siblings: All	0.007 (0.004)	0.005 (0.005)	0.009 (0.005)
Wealth Proportion	0.009 (0.008)	0.010 (0.008)	0.009 (0.011)
Percentage Any Secondary School: All	-0.009 (0.012)		
Percentage Any Secondary School: Boys		-0.018 [†] (0.010)	
Percentage Any Secondary School: Girls			0.008 (0.016)
Year Linear Trend	0.001* (0.000)	0.001 [†] (0.000)	0.000 (0.000)
Constant	0.368 [†] (0.183)	0.324 [†] (0.167)	0.837** (0.305)
Mean of Dependent Variable	-0.04	-0.03	-0.05
Number of Observations	48	48	48
R^2	.45	.43	.48

Notes: Robust standard errors are shown in parentheses. The dependent variable is the estimated country-year-specific two-stage least squares (2SLS) coefficient on the number of siblings predicting the probability of the child having any secondary school, using the sex composition of the first two births in the family as instruments. See an example of the 2SLS models run for Cambodia in Tables S3 (first stage) and S4 (second stage) in Online Resource 1. Regressions are weighted by the inverse of the standard error of the dependent variable estimated in the micro-level 2SLS models. All covariates at the country-year-level are calculated for the full sample in all columns except percentage any secondary school, which is calculated for the corresponding sample (both genders, boys only, and girls only) in each column.

[†] $p < .10$; * $p < .05$; ** $p < .01$

Hausman tests failed to reject the consistency of the more-efficient WLS models; thus, we report only WLS results and not the less-precise FE models. Among the four explanatory country-year-level variables, the only one that appears to be consistently statistically significant across samples is the percentage of last births that are male. The coefficient on the full sample (column 1) suggests that a 1 percentage point increase in male last births is associated with a 0.86 percentage point decrease in the magnitude of the IV coefficient. This is a nontrivial effect given that it is approximately one-fifth of the mean IV coefficient across country-years. Although the results for boys (column 2) are similar to those of the full sample, the effect of proportion male last birth on IV coefficients estimated on girls (column 3) more than doubles and is more than one-quarter of the mean IV coefficient across country-years for girls.

By contrast, almost all the other explanatory variables have coefficients that are very small and statistically indistinguishable from 0. The only exception is the coefficient on the percentage with any secondary school among boys, which is negative and marginally significant. Thus, in a setting where a high percentage of boys have secondary schooling, Q-Q trade-off for boys tends to be somewhat more prominent. In addition, the coefficient on the year linear trend is also significant for the full sample and for the boys-only sample. This finding suggests that the trend of Q-Q trade-off decreased slightly over time, which is somewhat expected given the improvement in life quality worldwide in recent decades.

One potential concern with the regression-estimated relationship between the IV coefficient and percentage of last births that are male is that the significance of the regression coefficient could be driven by only a few countries. Vietnam and Turkey are two outlier countries, with both high percentage of last births male (more than 0.55 in almost all years) and relatively large magnitude of IV coefficients. We thus reestimate the WLS model after dropping all observations from Vietnam and Turkey. The results are qualitatively similar to those presented in Table 3. The coefficients on percentage of last births male remain statistically significant, with a larger magnitude for the full sample (-1.141 , $p < .01$) and for the boys sample (-1.309 , $p < .01$), and smaller and insignificant magnitude for the girls sample (-0.937 , $p = .14$). However, the coefficients on number of siblings become marginally significant for the full sample and the girls sample ($p < .10$) after Vietnam and Turkey are dropped from the regressions; nevertheless, they are still within the confidence interval of the original coefficient including those two countries.

For comparison, we also estimate similar WLS models in predicting the micro-level OLS coefficients. The results are presented in Online Resource 1, Table S5. In this set of regressions, mean number of siblings is the only country-year characteristic with statistically significant coefficients: a one-unit increase in the mean number of siblings is associated with a 1.8 percentage point increase (decrease in absolute value) in the OLS coefficient for the full sample and a 2.2 percentage point increase for girls. The same coefficient for the boys sample is smaller and insignificant. The coefficients on the year linear trend are also positive and statistically significant, again for both the full sample and the girls sample.

Discussion

Results from 48 country-years examined in this study provide overwhelming evidence of a negative association between the number of siblings of a child and his or her

educational attainment. Children in larger families are less likely to have any secondary education in a variety of countries in Asia and Latin America and for several periods during recent decades. This association is present in plain OLS models and IV regressions using gender composition as instruments. One additional sibling in the family reduces the probability of secondary education by an average of 6 percentage points for girls and 4 percentage points for boys. In the context of recent literature about demographic dividends (Cuaresma et al. 2014), these results suggest that families do use the dividend to improve the education of children.

Returning to the questions posed at the outset of this article, we find a small decreasing trend over time in the size of the estimated Q-Q trade-off in the full sample and the boys sample. Results do vary somewhat by country: for example, in Costa Rica, the one prior country in which effects were examined over time, Li et al. (2014) found that the effects diminished substantially among girls, but they found a change in Q-Q trade-off for boys that is extremely consistent with the time trend found in the current study.

One pattern that was apparent in almost all the countries examined was that Q-Q trade-offs were substantially larger for girls than for boys. If interpreted causally, this would suggest that girls' secondary education attainment benefits more strongly from fertility decline than does boys'. To further explore this phenomenon, Fig. 5 plots the predicted educational increase per average decade of fertility decline, using each country's IV coefficient. The figure illustrates that indeed girls in several countries could have benefitted substantially more than boys. (Point estimates appear unbelievably large in Turkey and Vietnam but are within more credible ranges in several other countries.) In many other countries, however, the fertility decline can statistically explain only a small portion of the actual increase in educational attainment.

We investigated several country-year characteristics to help better explain these varying Q-Q effects across countries. The one that was most strongly related to the pattern of IV results was son preference, as operationalized by the percentage of last births that were male. This measure showed strong and increasing son preference in Asian countries in our sample, but it was weak and unchanging in Latin America. In

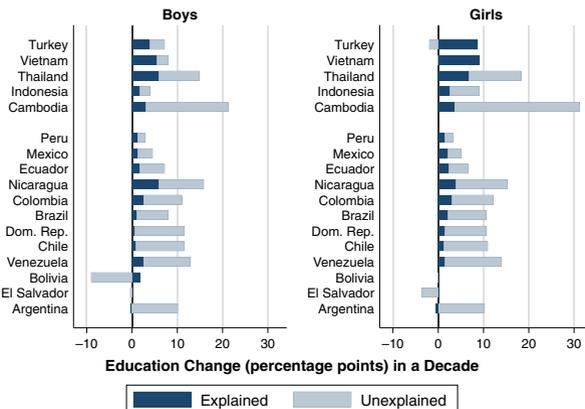


Fig. 5 Education increase attributable to fertility decline. Education change “explained” is the change in education predicted per decade with the IV model. Countries are sorted by both continent and proportion explained among girls

both regions, though, stronger son preference was related to stronger Q-Q trade-offs (as estimated by the IV models, which we believe are more reliable than the OLS results); this was true for both boys and girls, although the relationship was even stronger for girls.

We considered several explanations for the strong correlation between son preference and Q-Q trade-offs across country-years. Families with strong son preferences may also place less value in children's education in general, which affects their education investment decisions. In other words, in a cultural or social context where son preference is prevalent, parents are more likely to sacrifice children's education when faced with possible resource constraints arising from having an additional child. Further, when this occurs, girls in such families will be harmed more than boys because the parents will prioritize sons' education over daughters' (again, because of son preference). Note that son preference can be prevalent even in a country with high *overall* level of secondary school enrollment, such as Vietnam: the high overall level of secondary school enrollment largely reflects the quality of the government institution, which can mask variation in education investment across individual households due to Q-Q trade-off. Therefore, this explanation does not contradict the fact that we control for secondary education enrollment in the macro-level regressions.

The relationship between the preceding result and the local average treatment (LATE) interpretation of the IV coefficients is worthy of further discussion. LATE suggests that the IV coefficient is a weighted average of the treatment effects among those observations whose treatment is actually affected by the instruments (compliers). Because we use the sex composition of the first two births as instruments, a potential concern is that son preference strongly predicts higher Q-Q trade-off simply as a mechanical result of LATE—that is, that only households with son preference in a given country-year contribute to the estimation of the IV coefficients. However, as evidenced in Table S2 (Online Resource 1), many countries—mostly in Latin America—exhibit preferences for balanced gender among children. As such, families with balanced gender preference for children will also contribute to the estimation of the IV coefficients, which makes it very unlikely that the highly significant coefficients on last birth male in the macro-level regressions are simply an artifact of LATE.

Overall, estimated Q-Q trade-offs were larger in the Asian countries examined than in Latin American countries, especially for girls. Interestingly, the IV estimates of Q-Q trade-off magnitudes did not appear to vary systematically with the overall level of fertility (average number of siblings) or development (percentage attending secondary school, or the crude census wealth measure). Thus, among the variables examined, the lower son preference in Latin American countries is the only one that significantly helps explain the smaller Q-Q trade-offs compared with Asia. Given this, the R^2 (approximately .5) in the macro country-level WLS regression is perhaps surprisingly high; this could raise multicollinearity concerns, but further testing indicates that adding each of the variables individually to the random effects regression leads to similar results, as displayed in Table 3.

Several other methodological issues should also be considered when interpreting results. Using sex composition as instruments in studying Q-Q trade-off is subject to several limitations according to existing literature (e.g., Åslund and Grönqvist 2010; Shultz 2007). Li et al. (2014) discussed two main broad critiques of this instrument. First, the gender of a child can be affected by mother's nutrition because low nutrition

intake is associated with a lower probability of male births (Almond et al. 2011; Mathews et al. 2008). Second, gender composition of siblings may affect child outcomes through mechanism other than sibling size (Åslund and Grönqvist 2010).

Malnutrition due to fasting during Ramadan, as examined in Almond et al. (2011), is not likely to be an issue in Asian and Latin American countries, where no such tradition exists. Moreover, the smaller effect size that Mathews et al. (2008) found suggests that any confounding effect is unlikely to be significant. On the other hand, an alternative mechanism through which sibling gender composition affects children's education is mothers' marital status. Using data from the Current Population Survey, Dahl and Moretti (2008) found that women with firstborn daughters are less likely to marry but are more likely to be divorced if they had ever married. Unfortunately, our data do not allow us to convincingly rule out either possibility. However, we test them indirectly by examining the gender of the first and second births by wealth level in one particular country: Vietnam. We choose Vietnam because it has both relatively low gross domestic product (GDP) per capita and strong son preference as demonstrated from our previous results. If anything, the preceding two scenarios should be more prominent in Vietnam than other countries. Contrary to what the aforementioned critiques may predict, we find that families with asset index below the median are actually less likely than those above the median to have a first born daughter or to have both first- and second-born daughters.

On the other hand, one explanation for this result could be sex-selective abortion: that is, families with stronger son preference or lower socioeconomic status may be more likely to choose abortion when they find out during mother's pregnancy that the fetus is female. Evidence for this phenomenon in Vietnam does exist (Bélanger 2002). Predicting the direction of bias with sex-selective abortion is more difficult because it affects both the first-stage and the reduced-form coefficients simultaneously: that is, families who choose sex-selective abortion are more likely to have higher fertility as well as have children with lower educational attainment. Therefore, had there been no sex-selective abortions, the dummy variable for first two girls would have coefficients with smaller magnitude in both our first-stage (with negative sign) and reduced-form regressions. Because the IV coefficient equals the reduced-form coefficient divided by the first-stage coefficient, the direction of change in the IV coefficient is ambiguous when there are decreases in both the numerator and denominator. Nevertheless, the fact that omitting Vietnam and Turkey does not significantly change the results is somewhat reassuring.

Another possible mechanism through which sex composition may affect children's education is the reference group effect proposed by Butcher and Case (1994), who found that women raised with only brothers received, on average, significantly more education than women raised with any sisters, controlling for household size. Their findings are consistent with the reference group model, which suggests that the presence of a second daughter in the household changes the reference group for the first because parents with only one daughter may measure her achievement on the same scale as their sons' and may provide her with an equal share of the household's educational resources as her brothers (Butcher and Case 1994). Relatedly, Morduch and Garg (1998) also found that children who had all sisters had better health outcomes than if they had all brothers—an effect that does not differ significantly by gender. To mitigate this in the IV model, we controlled for whether the child has any sisters, but it remains an area of methodological controversy.

We also look forward to additional future research conducting similar cross-national comparisons using other methodological approaches, such as using fertility variation arising from twins. More substantively, we encourage future work to understand estimated effects in those countries for which we have newly uncovered unusually large relationships (such as Turkey and Vietnam) as well as further in-depth exploration of other countries and hypotheses explaining cross-national variation. The major contribution of the present study is to establish estimated macro-level patterns across a wide range of countries; progress in this area will necessarily require an iterative process similar to ours in order to better understand individual countries and carefully chosen comparisons of pairs of similar or contrasting settings.

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