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Stacey Chen
Yen-Chien Chen
Jin-Tan Liu

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THE IMPACT OF FAMILY COMPOSITION ON EDUCATIONAL ACHIEVEMENT

By Stacey H. Chen, Yen-Chien Chen and Jin-Tan Liu*
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Parents preferring sons tend to go on to have more children until one or more boys are born, and to concentrate investment in boys for a given sibsize. Therefore, having a brother may affect child outcomes in two ways: indirectly, by decreasing sibsize, and directly, where sibsize remains constant. We develop an identification strategy that allows us to separate these two effects. We then apply this to capture the heterogeneous effects of male siblings in both direct and indirect channels, using 0.8 million Taiwanese first-borns. Our empirical evidence indicates that neither effect is important in explaining first-born boys' education levels. In contrast, both effects for first-born girls are evident but go in opposite directions, resulting in a near-zero total effect which has previously been a measure of gender bias. These results offer new evidence of sibling rivalry and gender bias in family settings that has not been detected in the literature.
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We find new empirical evidence of gender bias in family settings using unique data from Taiwan, an economy with a long tradition of son preference. As in most countries in East and South Asia, many parents in Taiwan follow a son-preferring fertility-stopping rule.[1] They tend to go on to have more children until a boy is born, and to concentrate investment in boys conditional on sibsize. Thus, having a brother may affect his sibling's human capital formation in two ways: indirectly, by decreasing sibsize (Das 1987, Jensen 2003) and directly, where sibsize remains constant as the family goes through adjustments.[2] Perhaps surprisingly, little

* S. H. Chen: Academia Sinica, chens@nber.org. Y. C. Chen: National Chi-Nan University. J. T. Liu: National Taiwan University and NBER. This is a heavily revised version of Chen, Chen, Liu and Lien (2009). Earlier versions of this paper were circulated under the title "We Prefer Sons, but Does It Matter? Evidence from Matched Administrative Data from Taiwan" and "Estimating the Causal Effects of Sibling Sex Composition on Child Mortality and Education Using Twin Gender Shocks." Thanks to the Ministries of Education and Interior Affairs for providing administrative data, and to Ming-Ching Luoh for providing tax data. We benefited from feedback from Josh Angrist, Esther Duflo, Nancy Qian, Francis Vella, Taylor VanderWeele, and participants in universities and the 2009 NBER Education and Children's Programs. We acknowledge financial support from the National Science Council (NSC 101-2628-H-001-001-MY3) and the National Health Research Institute.

[1]See Deuchler (1992), Das Gupta and Li (1999), and Croll (2000) for the literature on family systems and son preferences in East and South Asia.

[2]Given sibsize, a son's birth may induce parents to adjust their saving, relationships, labour supply, or resource allocation (Parish and Willis 1993, Garg and Morduch 1998, Morduch 2000, Lundberg and Rose 2002, Rose 2003, Dahl and Moretti 2008, Ananat and Michaels 2008, Wei and Zhang 2011). Jayachandran

empirical work has been done to study the relative importance of these two effects and the mechanism of how family composition affects child outcomes.

Identification of the direct effect of male siblings is particularly challenging because observed sibsize cannot be truly fixed with changes in sibling sex composition, as many families follow the son-preferring fertility-stopping rule.[3] If the direct effect is measured by fixing the observed sibsize, thus muting the indirect effect, then identification is restricted only to families who have no pro-male bias or do not follow the son-preferring fertility-stopping rule. As a result, either effect is understated or not well defined. If the direct and indirect effects go in opposite directions, then the total effect (that is the sum of both effects) understates the degree of gender bias even more.

A unique aspect of this work is separate identification of the direct and indirect causal channels by which sibling sex composition may affect the outcomes of older siblings,[4] as illustrated in Figure 1. We suggest that potential sibsize given sibling gender can be fixed in counterfactual worlds (where sibling gender is viewed as an assigned experiment), although in reality, observed sibsize cannot.[5] Motivated by VanderWheele's (2013) causal inference model, we redefine and estimate the direct and indirect effects of sibling gender on educational achievement using Rubin's (1974) counterfactual notation. To capture the causal chain from sibling gender to fertility choice, and eventually to child outcomes, we adopt simultaneous equations with an interaction between sibsize and sibling gender. The interaction term is required to avoid muting the indirect effect, so as to avoid the ill-posed definitional issue. As far as we know, we are the first in the literature to apply this strategy to detect gender bias in educational attainment.

This contribution is made possible by overcoming data limitations. Previously, child outcomes could be observed only during infancy, so there are few studies of parental decisions to invest in educating daughters versus sons. By linking the Birth Registry records of first-born babies between 1978 and 1984 to their College Entrance Test records at age 18, we directly observe child educational attainment of more than 0.8 million first-born children during their adolescence.

The key to separate identification of the direct and indirect effects is the unconfoundedness condition of child gender (that is, child gender is selected on observable factors, including a comprehensive list of family backgrounds and parental characteristics). Our data is from a population born prior to legalization of abortion and prior to prevalence of ultrasound technology (which can be used for

and Kuziemko (2011) offer important insights for the mechanisms of how sibling sex composition affects child health differently, before and after reaching the desired sibsize, via the mother's decision as to when to wean her child. In contrast, having a son in the U.S. may help other siblings to develop assertive attitudes toward greater success in attaining education (Butcher and Case 1994, Kaestner 1997).

[3]Oaxaca's decomposition cannot work when the grouping variable (that is sibsize in this paper) is a mediating variable which affects child outcomes and changes with sibling sex composition.

[4]We focus on the sibling-gender effect on older siblings, not on children born later, since fertility is endogenous to the sex of the previously born. Estimates of effects on later-born children would be difficult to interpret.

[5]Potential sibsize differs from the desirable sibsize if birth control has not been made available or if the mother is too old to conceive.

prenatal testing for child sex). Using sex ratio statistics and hypothesis testing, we show that endogeneity of child gender or sex-selective abortion is not a concern in our data, supporting the validity of the unconfoundedness condition. Additionally, endogeneity of fertility choice is adjusted by using the incidence of twinning at the second birth, as in Rosenzweig and Wolpin (1980), Black, Devereux, and Salvanes (2005), and Angrist, Lavy, and Schlosser (2010). We address the issue of potential endowment deficits in twins by fixing the initial health of the second-born, as recommended by Rosenzweig and Zhang (2009). One important side product of our analysis is consistent estimates of the family-size effect that vary considerably across sibling sex compositions. First-born girls with a younger sister receive a particularly large family-size effect, three times as great as that for those with a younger brother or for first-born boys regardless of sibling gender.

We provide strong evidence showing that our empirical strategy makes important differences. First, we establish evidence of extraordinarily strong demand for sons in Taiwan, seemingly contrary to the near-zero total effects of having a younger son on the first child's educational attainment. By applying our decomposition method, we find that both the direct and indirect effects are near zero for first-born boys. In contrast, first-born girls receive a negative direct effect and a positive indirect effect, both of which are large and significant (about 8 to 10 percent of the high school completion rate or university enrollment rate) and almost cancel each other out, the result being a near-zero total effect. The indirect effect is large and significant for first-born girls but has almost no impact on first-born boys. This is because having a younger brother considerably reduces potential sibsize of a first-born girl, but has a much smaller impact on that of a first-born boy. This phenomenon cannot be discovered without interacting sibsize with sibling gender in the system of simultaneous equations.

Second, the adjustment for endogenous sibsize is large and significant in estimating the sibling-gender effects and the family-size effect. Contrary to near-zero sibling-gender effects based on Ordinary Least Square (OLS) methods, the Instrumental-Variable (IV) results show considerably large direct and indirect effects of having a younger brother on first-born girls' educational achievement on average. Additionally, the difference between OLS and IV results gets much larger if the interaction term is included, suggesting that the sibling-gender effects change with sibsize. Indeed, inclusion of an instrumented interaction term more than doubles the estimated direct and indirect effects of sibling gender on first-born girls, but has little impact on first-born boys. These findings highlight the importance of adjusting for endogenous sibsize in estimating the degree of gender bias in family settings.

The remainder of this paper is organized as follows: Section I introduces our data sets, reports descriptive statistics, and examines endogeneity of child gender. Section II states the concepts and notations used throughout the paper and describes our empirical strategies. Main empirical findings are summarized in Section III. Section IV concludes.

I. Data and Background

Identifying the causal impact of a change in sibling sex composition on educational achievement requires a large amount of detailed data. The data should contain information about sibling sex composition and child educational attainment up to at least the late teens. To fulfill this requirement, we link two national administrative datasets of Taiwan, namely, Birth Registry records and University Entrance Test records, using children's unique ID numbers. To control for birth-order effects, we focus on education of the first-born children from families with at least two children.

Our master data file is the Birth Registry of Taiwan since 1978, the initial year of the digitization of the data. It contains information on each newborn child's birth weight and birthplace, parents' education, and everyone's birth date and unique identification number. We focus on the sample of first-born singletons, born prior to January 1, 1985, when the Eugenics Protection Law began to be enforced. Although prenatal sex-testing by ultrasound began during the early 1980s, it was only after 1986 that the technology for sex testing became widely available; however, it remained limited to singletons of higher birth orders (Lin, Liu and Qian 2014).

In extracting data on first-borns, we first identify mothers who had their first child during the years 1978 to 1984 at age 18 or older, and then determine whether the mother had at least two children. To accurately measure sibling sex composition, we trace all siblings of the first-born child by their mother's unique identification number for 15 to 22 years, up to 1999. No mother in our data had a child in either 1998 or 1999, so the measure of sibling sex composition is accurate. Taiwan has no birth-control policy in place, so our data are not distorted by under-reported female births induced by forceful birth-control policies.

Birth Registry has detailed categorical information about parental education. Because the number of years of education in the general versus vocational tracks are not truly comparable, we capture the variation in parental education by using five educational indicators: university degree or above, professional training degree, high school diploma, vocational high school diploma, and junior high school diploma. The excluded category, primary school or below, is the reference group.

Our analysis includes a total of 82,631 first-born singletons from the Birth Registry who have at least one sibling and have complete information about parental age and education, birth year and birthplace. First-born children are classified as receiving an intervention if the next sibling is a brother rather than a sister. In the control group, the next sibling is a sister. If the second birth results in mixed-sex twins or triplets, then we randomize sibling gender by the fraction of boys in the birth. We assign 1 to sibling gender with probability of a boy in mixed-sex twins (0.5), the probability of a boy in triplets with one boy (0.33), and the probability of a boy in triplets with two boys (0.66); otherwise, we assign 0. In the full sample, 424,166 first-born children are in the treatment group having a second-born brother, whereas 397,465 are in the control group having a second-born sister.

We acquired education information from University Entrance Test records of 1996 to 2003 when the first-born just turned 18. The data included two sets of tests: the general tests, conducted in February during the high school senior year, and the union entrance tests, conducted in July after high school graduation. These tests offer two distinct channels to university education: students can apply for university admissions using their general test scores and skip the tests in July. If their application results are not satisfactory, students can forego prior admissions and take the union entrance tests in July after graduation. The indicator for university admission in our study is based on both tests, or channels. In addition, we construct an indicator for high-school completion using "took general tests in February" as a proxy because most graduating seniors take the tests. It is noteworthy that our calculation of high school completion and university attainment excludes vocational high school and vocational college education.

Since 1928, when the first university was founded, the brightest students in Taiwan have attended public universities. During our sample years of 1996 to 2003, tuition and fees in public universities were about 14 percent of yearly family income, whereas the cost of attending private colleges was about 25 percent. Additionally, we control for family socio-economic status by linking our first-born sample to data on per capita taxable income in district of birth, provided by the Ministry of Finance.

A. Characteristics of the First-Born

Table 1 reports the statistics of first-born outcomes and characteristics by sibling sex composition. About a quarter of first-born children completed high school, and only 15 to 18 percent enrolled in universities. First-born girls have larger families and they are about 2.5 percentage points more likely to enroll in universities than their male counterparts. First-born child education does not seem to change with the gender of the next child, but sibsize varies drastically with sibling sex composition. Families with two girls have 0.54 more children than those with two boys and 0.43 to 0.44 more children than those with a mixed-sex composition. Unlike American parents who prefer a mixed-sex composition, Taiwanese parents strongly favor multiple sons.

In our data of the first-born children, mothers with two girls are 28 percentage points more likely to have a third child than those with two boys. That differential is less than two percentage points in the United States and Israel (Ben-Porath and Welch 1976, Angrist and Evans 1998, Angrist, Lavy and Schlosser 2010).

B. Demand for Multiple Sons

Taiwanese have a long tradition of pro-male bias owing to cultural factors. Confucianism – the grounding philosophy in Taiwan, Japan, Korea, and imperial China – dictates social statutes and provides rationales for the subordination of women to men, within a strict family hierarchy. According to Confucianism, family line and wealth should be transmitted from father to son, irrespective of ability, except in cases where there is no direct male line. In return, sons and their spouses assume responsibility for taking care of the parents if they are too infirm to work. In contrast, daughters move out of the family household at the time of marriage. These social norms have been acting as old-age social security for the elderly for centuries, in the form of extended families, composed of sons (and their spouses, if married), unmarried daughters, parents, and grandparents. Although government-funded old-age social security in Taiwan began during recent decades, the extended families (even if they do not live together) are still the primary source of support for the elderly. Thus, the demand for old-age social security is more likely to be met by having more sons.

The Confucian thought and discipline that systematically justifies the demand for multiple sons, such as *Analects* (ca. 479 BCE), was at the core of the educational curriculum in Imperial China for more than two millennium.[6] It remains a dominant component of the educational curriculum in Taiwan.

Using the Taiwanese Birth Registry, we measure the demand for sons by the effect of a change in sibling sex composition on sibsize, conditional on observed family backgrounds. Table 2 suggests that Taiwanese families strongly prefer sons to daughters, and multiple sons to mixed-sex composition, and the tendency gets stronger if the mother is less educated or if the child was born in a rural area. Model (I) in the top panel shows that having a son, regardless of the birth order, decreases sibsize by 0.43 to 0.44 person. Because birth order is unimportant in explaining the demand for sons, we use Model (II) where we focus on the impact of sibling sex composition on sibsize, leaving out the factor of birth order. The results suggest that, compared to families with two boys, those with two girls have about 0.53 more children, and those with mixed-sex composition have about 0.1 more. These estimates are extraordinarily large, since they account for approximately 20 percent and 4 percent, respectively, of the average sibsize (about 2.7). If the child was born in urban or the mother has a junior high school diploma or above, then the level of these estimates is reduced by about one third. These results are robust and precise, regardless of whether or not we include parental education, per capita taxable income in the district of residence, or both.

C. Testing for Exogeneity of Sibling Sex Composition

Although the presence of sex-selective abortion is neither observable nor testable in the data, we examine exogeneity of sibling sex composition in four ways. First,

[6] A Chinese poem, dating from centuries before Confucius, "Si Gan" from *Book of Songs* (or *Shi-Jing*), which is believed to have been compiled by Confucius, advised parents to allocate family resources unevenly between sons and daughters: "When a son is born, let him sleep on the bed, dress him with fine robes, and give him jade to play... When a daughter is born, let her sleep on the ground, cover her in usual wrappings, and give her tiles for playing." Perhaps this is the oldest text on gender bias.

the ratio of boys to girls at birth is approximately 1.044 for first-borns, and 1.067 for second-borns. Both ratios are within the range that demographers consider normal on the basis of historical evidence (see, e.g., Johansson and Nygren 1991).

Second, we compare demographics of the full sample (born between 1978 and 1984) with the cohort born prior to 1980 when ultrasound (the technology for prenatal testing for child sex) was not yet available. As shown in the Appendix Table A1, the full sample and the pre-1980 birth cohort share similar demographic statistics, except that the full sample has considerably higher parental education owing to the introduction of nine years of compulsory education in 1968, which affected the parents of the younger cohorts. In Column (3), we further restrict the sample to those whose next sibling was born prior to 1985, the year when abortion was legalized. This restriction has little impact on the sex ratios. The sex ratios, 1.044 and 1.067, of the full sample are less male-dominated than those of the pre-1980 cohorts, which had no sex-testing technology available. Although some second-born children in our data were born after 1985 and might have been exposed to ultrasound, we still include them in the data so we do not restrict our analysis to the families with shorter birth spacing, who might have a stronger demand for sons (see, e.g., Jayachandran and Kuziemko 2011).

Third, the R-squared adjusted in Table A2 for a regression of $Boy2nd$ (the sex of the second-born) on $Boy1st$ (the sex of the first-born) and a comprehensive list of observed family backgrounds is close to zero, and the implied F statistic is below the critical value at the 99 percent significance level. We note that having a first-born girl is associated with a 0.33 percentage point increase in the ratio of boys to girls at the second birth and it is statistically significant. Nevertheless, the sex ratio at the second birth after accounting for this addition (1.067+0.0033) remains within the normal range.

Finally, regressions of birth spacing between the first two children on sibling sex composition and family backgrounds provide no evidence that birth spacing is distorted by the period of time over which a female fetus is conceived and aborted. Table A3 shows that the estimated coefficient of the interaction between $Girl1st$ and $Boy2nd$ is only four days and it was statistically insignificant, so we reject the hypothesis that after having given birth to a girl first, the mother tends to spend more time trying to bear a boy relative to a girl. Our statistical results suggest that sex-selective abortion is not a concern among children of the first two births in our data.

II. Empirical Strategies

In this section, we discuss the conceptual basis of our empirical analysis, beginning with how direct and indirect effects can be defined and separated. The key is that potential sibsize can be fixed in counterfactual worlds, although observed sibsize cannot. Under the unconfoundedness condition, we show that the conventional measure for the direct effect (denoted by CDE) is biased downward and the bias is proportional to the distance between the unconditional versus conditional mean of sibsize, given sibling gender. Our entire analysis is conditioned implicitly on a set of covariates, X, which includes the gender of the first-born child and family background variables listed in Table 3.

Considering the randomized gender of the second-born sibling D affects sibsize M and the first child's outcome Y. We use Y_{DM} to denote the potential outcome of the first child given sibling gender D and sibsize M; we use M_D and Y_D to indicate the potential sibsize and the potential outcome given the gender of the next sibling D. We temporarily assume that we observe $(M_0, M_1) = (m_0, m_1)$ for each first-born child in the data. We define:

- **Controlled direct effect** $CDE = Y_{1m} - Y_{0m}$;
- **Direct effect** $DE = Y_{1m_1} - Y_{0m_1}$;
- **Indirect effect** $IE = Y_{0m_1} - Y_{0m_0}$; and
- **Total effect** $TE = DE + IE = Y_{1m_1} - Y_{0m_0} = Y_1 - Y_0$.

The conventional measure of the direct effect is CDE, defined by fixing observed sibsize and by assuming that sibsize cannot change with sibling gender, $m_0 = m_1 = m$. This forces the implied indirect effect to be zero so CDE is actually a conditional total effect, restricted only to families with no gender bias in fertility choice. We propose to measure the direct effect by DE conditioning on potential sibsize m_1 as though every family had a male second birth.

The direct effect (CDE or DE) contains all sources by which sibling gender may affect the first child's outcome, not via a change in sibsize. These include (part of) the changes in parents' savings, consumption, relationship, labor supply, and time allocation, as listed in footnote 2. In contrast, the indirect effect, IE, depends entirely on a change in parental fertility choice, in response to a change in the sex composition of previous children. A stronger demand for sons induces a larger indirect effect. Thus, the indirect effect is not negligible for regions where the son-preferring fertility-stopping rule is common.

We cannot identify (DE, IE) because they depend on potential sibsizes m_1 and m_0, which cannot be both observed for the same person. Our target parameters are the average direct and indirect effects, denoted by (ADE, AIE), which can be constructed by averaging the DE and IE over all possible values of potential sibsize in $\mathcal{M}$.

$$\text{(1)} \quad ADE = \sum_{m \in \mathcal{M}} E[Y_{1m} - Y_{0m} | M = m] Pr\{M = m | D = 1\},$$
$$AIE = \sum_{m \in \mathcal{M}} E[Y_{0m} | M = m][Pr\{M = m | D = 1\} - Pr\{M = m | D = 0\}].$$

Mechanically, $ADE + AIE = ATE \equiv E[Y_1 - Y_0]$ so we can derive AIE simply by subtracting ADE from ATE. Here we can use the mass function of observed sibsize, $Pr\{M | D = d\}$, to capture the mass function of potential sibsize,

$Pr\{M_d|D=d\}$, because sibling gender D is randomized. We illustrate this idea using the case of binary fertility choice, $M = Morethan2 \in \mathcal{M} = \{0,1\}$; $M=1$ if parents have a third child and 0 otherwise. For binary fertility choice, we have:

$$Pr\{M=1|D=d\} = E[M|D=d] = E[M_d|D=d].$$

Additionally, we assume that the relationship between observed and potential outcomes satisfies the following equation:

$$E[Y|D=d, M=m] \quad = \quad E[Y_{dm}|D=d, M_d=m].$$

Given randomized sibling gender, each person's M_0 and M_1 are missing at random so we can use the conditional expectation of the observed outcome to impute the conditional expectation of the potential outcome.

Because fertility choice is endogenous, we instrument $Morethan2$ (the decision to have a third child) by Z, the twinning indicator at the second birth. To address the endowment deficit of twins, we include in the covariate set X, the initial health condition of the twins at the second birth, as suggested by Rosenzweig and Zhang (2009). We assume Z to be exogenous given X.

To estimate the target parameters ADE and AIE, we begin with a linear probability model with constant coefficients:

$$\begin{aligned} M &= \alpha_0 + \alpha_1 D + \alpha_2 Z + u, \\ Y &= \beta_0 + \beta_1 D + \beta_2 M + \epsilon. \end{aligned}$$

The Greek letters are coefficients. The outcome residual ϵ can be correlated with the selection error u. However, this model is too restrictive because it assumes that effect-heterogeneity is absent in both observables and unobservables. In particular, this model implies $ADE = DE = CDE = \beta_1$, where the controlled direct effect is seemingly a correct measure for direct effects but it is entirely driven by the functional-form assumption.[7]

To allow effect-heterogeneity in observables and unobservables, we consider a more flexible model. First, we add an interaction term, $D \times M$ in the outcome equation, and instrument the interaction by $D \times Z$. Second, we replace the outcome error term ϵ with ϵ_{DM} to allow for random coefficients.

$$\begin{aligned} M &= \alpha_0 + \alpha_1 D + \alpha_2 Z + \alpha_3 D \times Z + u, \\ Y &= \beta_0 + \beta_1 D + \beta_2 M + \beta_3 D \times M + \epsilon_{DM}. \end{aligned} \tag{2}$$

We assume ϵ_{0M} and ϵ_{1M} share the same distribution, and both are correlated

[7]As an anonymous referee has noted, in the case of constant-coefficient with no interaction, the IV estimate is likely to be outside of the convex hull of the local average treatment effect. This problem is a consequence of the ill-posed definitional issue in this specific model, not a direct result of instrumental-variable methods.

with fertility choice M through the selection error u. This model suggests that the average direct effect of sibling gender, defined in (1), can be rewritten as:

$$\begin{aligned}(3)\qquad ADE = {} & \beta_1 + \beta_3 E[M|D=1] \\ & + \sum_{m=0,1} E[\epsilon_{1m} - \epsilon_{0m}|M=m]Pr\{M=m|D=1\}.\end{aligned}$$

The third term is zero because ϵ_{0m} and ϵ_{1m} have the same conditional mean. If β_1 and β_3 are both identified, then we can identify the average direct effect ADE, and thus $AIE = ATE - ADE$ is also identified.

For comparison, the conventional measure for the average direct effect, $CDE = \beta_1 + \beta_3 E[M]$, can be estimated too. CDE is a biased measure for ADE unless fertility choice M is independent of the sex composition of the previous children, $E[M|D] = E[M]$. But if parents adopt son-preferring fertility-stopping rules, then $E[M|D=1] < E[M]$ and the bias of CDE is proportional to the difference between $E[M|D=1]$ and $E[M]$.

We estimate β_1 and β_3 (and thus ADE and AIE) by standard Instrumental-Variable methods using the twins instrument. The IV estimates only give us a causal interpretation that is limited only to complying families, whose sibsize would rise with twinning at the second birth, $M(Z=1) > M(Z=0)$. Consequently, estimation using all families, including those who would have had more than two children even in the absence of twins, may give different results from the IV estimates.

However, if the compliance rate is closer to one, then the IV estimates should be close to the average effects for the general population. Our first-stage estimates in Table 3 imply the compliance rate, given the sex of the next sibling D,

$$Pr\{M(Z=1) > M(Z=0)|D\} = E[M|Z=1, D] - E[M|Z=0, D],$$

ranging from 34 to 61 percent, when M indicates the choice of having a third child. This suggests that our IV results should come close to the population average treatment effects.

III. Results

Using the first-born data from families with at least two children, we generate three sets of results: first-stage estimates of endogenous fertility choice using twins at the second birth as an instrument, second-stage estimates of the human-capital-formation function of the first-born, and decomposition of the total effect into direct and indirect effects. We show that conventional methods (using either the total effect or the controlled direct effect) might have systematically understated the degree of gender bias in family settings.

A. First-Stage Estimates

In the first stage of estimation, we instrument the endogenous fertility choice of having a third child using twinning at the second birth. The first-stage results are very strong for both of the fertility-choice variables, $Morethan2$ and $Sibsize$, and their interaction with sibling gender $Boy2nd$. The estimates in the top panel of both Tables 3 and A5 suggest that a twin birth increases the probability of having more than two children by 34 to 60 percentage points. Those in the bottom panel show that a twin birth increases the completed sibsize by about 0.6 to 0.7 children. These estimates are all significant with small standard errors.

The first-stage estimate decreases with the number of boys as first births because of strong demand for (multiple) sons. If the first two births are both boys, twinning increases the fraction of families to have a third child by more than 60 percentage points. This number goes down to 55 percentage points if only one of the first two births is a boy. The number goes further down to 34 percentage points if they are all girls because some parents who have no sons keep trying, whether or not they give birth to twins. Compared with families with all girls, a family with twins is more likely to push parents with at least one boy above their optimal number of children, so their first-stage estimates tend to be greater.

While having a second-born son markedly drives up the effect of having twins on the desire to keep on trying, it has a small and insignificant impact on the completed sibsize. The estimated coefficients of the interaction term in the bottom row of Table 3 indicate that having a second-born boy increases the effect of twins on $Sibsize$ only by 0.021 or less, and it is statistically insignificant. Nevertheless, the coefficient of the interaction $Twin2nd \times Boy2nd$ is large and significant in the first-stage regression of the interaction term $Sibsize \times Boy2nd$, as Table A5 shows. The coefficient of the interaction is about 0.52 to 0.70, with very small standard deviation. These figures suggest that we need not be concerned about weak instruments.[8]

B. OLS and 2SLS Results

In Table 4, we compute the regression of the first-born's completion of high school with family-composition variables, including sibsize and sibling gender. We find that the OLS method considerably understates the family-size effect on the first-born girl's education, because of the omitted-variable bias and the omitted interaction term. The downward bias is smaller for first-born boys.

OLS-estimated coefficients of sibsize and sibling gender in columns (1) and (6) are all small and negative for first-born girls. However, it is likely that unobserved bias against daughters is greatest for girls from large families with at least one male sibling, so both coefficients for first-born girls may be understated.

[8]Though not reported here, the estimated first-stage impact of having a son at the second birth on fertility choice of "trying again" for another son gradually decreases with higher parities and eventually fades away as sibsize is complete.

To tackle the unobserved bias that may affect both sibsize and the first child's outcome, we instrument the fertility-choice variable using twinning at the second birth. Compared to the OLS results, the 2SLS estimated coefficient of $Boy2nd$ more than doubles for first-born girls as expected, and little is changed for first-born boys. These results can be found in columns (2) and (7) of Table 4.

As sibling gender is essentially randomly assigned in our data (recall Table 1), a large change in the coefficient of the random variable in column (2) after instrumenting fertility choice is noteworthy. The 2SLS estimated coefficient of sibsize for first-born girls also rises substantially. It is likely that, for first-born girls, the sibling-gender effect rises with sibsize, or that the sibsize effect rises with the presence of a brother. In either case, we should allow sibsize and sibling gender to interact in the regression analysis.

As expected, adding an interaction term between sibsize and sibling gender in columns (3) and (8) considerably changes the 2SLS result for first-born girls. The coefficient of the interaction term is large and significant, and the coefficient of $Boy2nd$ rises at least sevenfold. In contrast, both coefficients for first-born boys remain small and insignificant.

It is worth emphasizing that the coefficient of $Boy2nd$ on child outcomes generally cannot be interpreted as the direct effect of sibling gender on education, as noted in Section II. In contrast, the 2SLS estimated coefficient of fertility choice still has important causal interpretations for the impact of family size on child education. The result in column (3) indicates a clear tradeoff between child quality and quantity when there is no son at the first two births. The average high-school completion rate falls by at least 10 percentage points with more than two children in the family, or by 5.3 percentage points with one additional sibling. This is extremely large because they account for 40 percent and 20 percent of the high school completion rate, respectively. The largest family-size effect appears among first-born girls whose next sibling is a girl, because parents who would keep on trying after having all girls in the first two births are most likely to invest only in the later-born son, compared to those who stop.

By contrast, if there are one or more sons in the first two births, then the family-size effect is reduced slightly and becomes imprecise. As columns (3) and (8) suggest, having more than two children with at least one boy decreases the high school completion rate by 4 percentage points or less with standard errors being around 2 to 3 percentage points. Although these estimates are large, they are too imprecise to be conclusive.

C. Decomposition Results: The Average Direct and Indirect Effects

With the extremely strong demand for sons, it is perhaps surprising that Taiwanese girls on average are more likely than boys to complete high school and enroll in university. And the average total effect of having a second-born boy on whether the first-born girl completes high school is positive or nearly zero, as the first row of Table 5 shows. In contrast, the same effect is negative for first-born boys. These statistics might be seen as evidence for the absence of rivalry effects of male siblings on Taiwanese girls, even with exceedingly strong demand for sons. The key to explaining this puzzle is the presence of positive indirect effects owing to son-preferring fertility-stopping rules. Reduced sibsize after having a subsequent brother allows more of family resources to be invested in the first-born girl's education. Since indirect effects run in the opposite direction from direct effects, the average total effect is close to zero. We enlarge on these results below.

Columns (1) to (5) of Table 5 report our decomposition results for first-born girls. The estimated average direct and indirect effects rise considerably from column (1) to column (2), after we address the endogeneity of fertility choice, suggesting substantial adjustments for endogenous sibsize in explaining the first child's education. The adjustments go further more after we add in column (3) an interaction between sibling gender and sibsize. This suggests a great deal of heterogeneity in the sibling-gender effects across various sibsizes. The interaction term should not be omitted from the model.

Unlike the large adjustments for endogenous sibsize among first-born girls, these adjustments among first-born boys are small and insignificant, as columns (6) and (7) show. This is because first-born boys have smaller families than first-born girls, regardless of whether their parents opt for child quality over quantity. In addition, first-born boys also have a much smaller adjustment for effect-heterogeneity across sibsizes, as columns (6) and (7) indicate. After having the first-born son, parental fertility choice or allocation of family resources do not seem to respond to the gender of the next sibling. As the average direct and indirect effects of sibling gender are both nearly zero, interacting sibsize with sibling gender has almost no impact on the estimation results.

We note that the gap in estimates between the controlled direct effect and the average direct effect is nearly zero for first-born boys, while it is much wider for first-born girls. After we address endogenous sibsize in column (2) and add the interaction term in column (3), the 2SLS estimated controlled direct effect amounts to about 60 percent of the average direct effect for first-born girls, while these two estimates are virtually identical for first-born boys. This contrast is due to the fact that the controlled direct effect is evaluated at the unconditional average sibsize (under the assumption that sibsize does not change with sibling gender), while the average direct effect is evaluated at the average sibsize conditional on the second child's being a boy. Owing to son-preferring fertility-stopping rules, having a second-born son is less likely to induce parents of first-born girls to keep trying, compared with those of first-born boys. As a result, the conditional mean sibsize of first-born girls, conditional on having a second-born brother, is smaller than the unconditional mean. In contrast, for first-born boys, the conditional and unconditional means of sibsize are unequal, so the bias of the controlled direct effect is approximately zero.

Overall, sibling-gender effects are much smaller on first-born boys than on first-born girls. The average indirect (direct) effect of sibling gender on first-born girls' education is more than 10 (4) times of that on first-born boys' education. This evidence shows a very strong pro-male bias, much stronger than what the controlled direct effect has indicated, and the opposite of what the total effect has suggested.

It is worth noting in Table 5, that our decomposition results are robust, regardless of which fertility-choice measure (either $Morethan2$ or $Sibsize$) is adopted, as long as we include an interaction term in the model and instrument endogenous fertility choice.

Decomposition results for another important education outcome – university admission at age 18 – show similar patterns, as column (2) of the top panel in Table 6 shows. The total effect of having a second-born brother on the first child's university enrolment is close to zero for both genders (the estimates for first-born boys are not shown), and it is large and positive for the first-born girls. On the basis of the estimates of the total effect, gender bias seems to be absent or at least not against girls. Only after we divide the total effect into direct and indirect effects does gender bias become evident. Unlike first-born boys, whose average direct and indirect effects are both close to zero, first-born girls receive a boost of 1.74 percentage points in their average indirect effect and suffer a loss of 1.54 percentage points in their average direct effect. Both estimates are statistically significant at the 99 percent level. These results are considerable in magnitude, since they account for about 10 percent of first-born girls' university enrolment rate, the same proportion as the effects on first-born girls' rate of high school completion, as reported earlier.

D. Exogeneity of the Twins Instrument

Although exogeneity of the twins instrument is not testable, we examine whether the occurrence of twins can be explained by parental education, place of birth, or the average taxable income in the district of residence. The F statistic (though not reported in tables) cannot reject the hypothesis that the coefficients for these background variables are jointly zero. This suggests that the birth of twins is not related directly to parental education and residential location. Another concern is that the occurrence of twins rises with maternal age at childbirth. We addressed this concern by including the age of the mother at the birth of her second child among our control variables. Our results change little with or without this covariate.

Exogeneity of the twin instrument has been questioned because twins have lower birth weight and shorter gestation duration than singletons. The subsequent birth of twin siblings likely has a direct effect on first-born children beyond just increasing sibsize. For example, compromised initial health of second-born twins may induce some parents to divert family resources from the twins to the first-born singleton (if parents have efficiency concerns), or the other way around (if parents have inequality aversion). In either case, the estimated family-size effect is biased, either downward or upward. Additionally, if parents preferring singletons are in favour of sons, the diversion of family resources might be greatest from girl twins to the first-born male singleton, thus the effects of having a brother on the first-born singleton will be understated, particularly among first-born boys.

To address the issue of endowment deficit of twins, Rosenzweig and Zhang (2009) suggest controlling for the initial health condition of the second birth, using their mean birth weight. The idea is that *fixing the birth weight of the second birth* in addition to observed family backgrounds, the only channel through which twinning at the second birth can affect the first-born child's education is via changing the sibsize. Their result, based on data from China, suggests that when mean birth weights are included, a second-birth twin pair negatively affects the outcome of second-born twins, but the twins' effect on the first child's outcome is small and insignificant, consistent with the assumption that the twins instrument is conditionally exogenous.

However, because boys are heavier than girls on average, part of the sibling-gender effect may be mistaken for birth-weight effects if we include the mean birth weight of the second birth as a control variable. Our results in column (5) of Table 5 show that adding the mean birth weight of the second birth leads to a 20 percent decrease in the 2SLS estimates of the sibling-gender effects among first-born girls. The downward adjustments for first-born boys in column (10) are smaller and imprecise. The 2SLS estimated family-size effects are also adjusted downward for first-born girls by 20 percent or more. Inclusion of the mean birth weight adjusts these 2SLS estimates downward because the birth weight may decrease with the occurrence of twins or the occurrence of a female singleton, either of which increases sibsize. We cannot truly fix the mean birth weight of the second birth when we estimate the sibling-gender effects. Thus, inclusion of the mean birth weight may open up another causal channel – from sibling gender, to the birth weight of the second-born sibling, and eventually to the first child's education. The mean birth weight becomes another mediating variable, like sibsize, in the model. Sibling gender may indirectly affect the first-born outcomes via changes in sibling birth weight, in addition to via changes in sibsize.[9]

One alternative control for the initial health of the second-born is the length of the gestation periods, which is not affected by gender. As statistics in Table 1 show, the gender gap in gestation duration is only 1 percent of the average duration and it is statistically insignificant. The result in column (3) of Table 3 suggests that one additional week in the second-born gestation period increases the likelihood of having a third child by 0.2 percentage points (SE=0.0005) if the first born is female. The second-stage result in column (3) of Table 4 indicates that the second-born gestation has almost no impact on the first child's education. The

[9]Unobserved confounding factors such as parents' lifestyle and characteristics are correlated with both birth weight and child outcomes. Birth weight too is likely to be endogenous, since it can be shaped by a wide range of factors, including maternal education, the introduction of social programs, and the interplay of genes and the environment. See Almond, Chay, and Lee (2005), Currie (2009), Almond and Currie (2011), and Currie (2011) for reviews of the literature. To formally estimate the indirect effect via changes in sibling birth weight, we need to instrument birth weight.

estimates based on gestation duration cannot support the conjecture that parents are in favour of singletons over twins due to their difference in initial health.

Irrespective of which of the initial health measures is used, the 2SLS estimated family-size effect for first-born girls remains striking (accounts for more than one-third of the high school completion rate) if the next sibling is also female. And the average direct and indirect effects of sibling gender still account for 8 to 10 percent of the average high school completion rate among first-born girls. In contrast, the family-size effect and the sibling-gender effect both remain nearly zero among first-born boys. Overall evidence strongly suggests that the roles of family size and sibling gender in forming human capital are much more important for first-born girls, compared to first-born boys.

E. Heterogeneous Effects of Sibling Sex Composition on Education

Gender bias among first-born children is most evident for urban families in term of entering university, especially if the mother gave birth in her early thirties. We report this result in Table 6 for first-born girls, as the estimates for first-born boys are all close to zero and insignificant.

In columns (5) and (6) of Table 6 where we focus on first-born girls in urban, we find that the average direct and indirect effects on university attainment are as high as 14 percent of the average enrolment rate in these areas, which is 40 percent stronger than the estimates in the full sample. In contrast, both effects on high school completion in the urban areas are proportional to those in the full sample. These results suggest that gender bias is equally strong in both urban and rural areas in terms of completing high school, but gender bias in terms of entering university is considerably stronger in urban than in rural areas.

In the bottom panel of Table 6, where we divide our first-born population by maternal age at birth, the most evident gender bias appears in attaining university education for first-born girls whose mothers gave birth in their early thirties. As column (4) shows, the average direct and indirect effects are both about 5 percentage points, in opposite directions, and both account for 15 to 18 percent of the enrolment rate. Compared to the results in columns (2) and (6), this is proportionally 50 percent stronger than the same effects in the full sample, and four times stronger than those with older mothers. Finally, if we restrict our first-born sample to those whose mothers have a junior high school diploma or above, both effects on first-born girls' education accounts for approximately 8 percent of the rate of completing high school or enrolling in university. This is proportionally similar to the results in the full sample, suggesting that the degree of gender bias is about the same across different levels of maternal education.

IV. Conclusion

Using an unusually large database from Taiwan, we first establish empirical evidence of an extraordinarily strong demand for sons. The size of a family having produced two daughters is 0.54 person larger than the size of a family having already produced two sons. Yet, contrary to popular beliefs for such regions where a preference for sons is prevalent, sample means show *no* evidence that boys have more opportunities than girls to complete high school or attend university. In fact, the total effect of having a brother, relative to a sister, is either positive or nearly zero on first-born girls.

We reconcile those seemingly contradictory results in two ways. We first allow sibsize to interact with sibling sex composition in the formation of human capital and then decompose the total effect of sibling gender on human capital into direct and indirect effects, with changes in potential sibsize as the mediator. We resolve the ill-posed definitions, caused by the fact that observed sibsize cannot be fixed with changes in sibling sex composition. We demonstrate that even if sibsize were exogenous, the ill-posed definitional issue exists. Additionally, we address the endogeneity of sibsize using the twins instrument conditional on various initial health measures.

Decomposition, with sibsize interacting with sibling sex composition, indeed makes important differences. We find that having a younger brother (as opposed to a younger sister) markedly lowers the potential sibsize of a first-born girl, to a degree that the positive indirect effect cancels out the negative direct effect on her education outcomes, resulting in a near-zero total effect. The positive indirect effect is driven essentially by external effects of having a younger brother, owing to a smaller family. Both the direct and indirect effects account for about one-eighth to one-tenth of the average educational achievement of first-born girls, while the effect of one additional sibling lowers her opportunity for a university education by about 10 percentage points (about two-fifths of the average university admission rate). Unlike the marked impact of sibsize and sibling gender on first-born girls, neither the number of siblings nor the gender of the next sibling has a noticeable effect on a first-born boy, regardless of the gender of the next sibling. This offers new evidence for gender bias in family settings that has not been reported in the literature.

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Figure 1. The Direct Effect (←) and the Indirect Effect (⇠) of Sibling Gender on Firstborn Outcomes

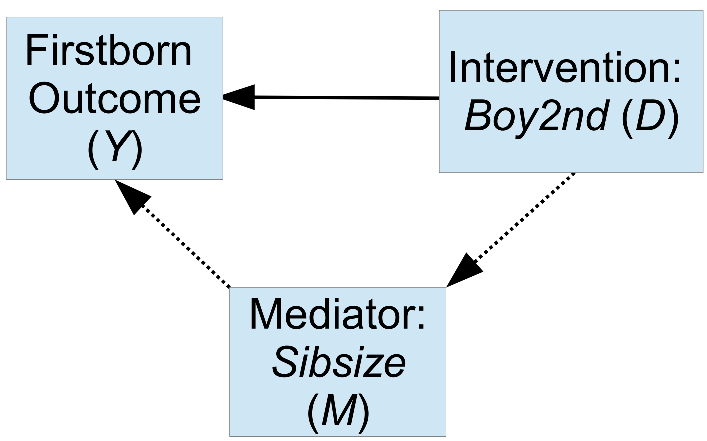

Table 1—Variable Means (Standard Deviations) for First-borns, by Sex Composition

	First-Born Girls		First-Born Boys	
	Boy2nd=0	Boy2nd=1	Boy2nd=0	Boy2nd=1
Sample size	193,731	208,169	203,734	215,997
Outcome variables Y				
High school completion	0.243	0.246	0.240	0.237
	(0.43)	(0.43)	(0.43)	(0.43)
Admitted to university	0.175	0.178	0.153	0.153
	(0.38)	(0.38)	(0.36)	(0.36)
Sibsize measures, mediator M				
More than two children $Morethan2$	0.707	0.484	0.490	0.425
	(0.45)	(0.50)	(0.50)	(0.49)
Complete $Sibsize$	3.046	2.606	2.612	2.509
	(0.91)	(0.73)	(0.73)	(0.68)
Instrument for fertility, $Z = Twin2nd$				
Twins at second birth	0.0071	0.0069	0.0067	0.0061
	(0.08)	(0.08)	(0.08)	(0.08)
Covariates X				
Mean birth weight of second birth (kg)	3.231	3.339	3.219	3.320
Gestation duration of second birth (weeks)	39.66	39.61	39.63	39.59
Urban (place of birth)	0.340	0.342	0.338	0.338
5-year average taxable income per capita in village (thousands)	730.0	730.8	729.6	729.2
Mother's age at second birth	26.2	26.2	26.2	26.2
Mother's year of birth	1957.3	1957.3	1957.3	1957.3
Father's year of birth	1954.0	1954.0	1954.1	1954.1
Mother's highest degree				
College/professional degree or +	0.070	0.072	0.071	0.070
High school diploma	0.063	0.063	0.062	0.063
Vocational high school	0.190	0.192	0.190	0.192
Junior high school	0.261	0.259	0.261	0.262
Father's highest degree				
College degree or above	0.064	0.065	0.064	0.064
Professional degree	0.075	0.075	0.075	0.075
High school diploma	0.094	0.095	0.093	0.094
Vocational high school	0.182	0.183	0.183	0.184
Junior high school	0.234	0.231	0.233	0.233

Table 2—Demand for Sons - Effect of Sibling Sex Composition on Sibsize

	All First-Borns			Mother Finished	Born in
Explanatory Variables	(1)	(2)	(3)	Junior HS or +	Urban
Sample size	821,631	821,631	821,631	266,463	278,934
Model (I)					
Boy1st	-0.432	-0.432	-0.432	-0.347	-0.366
	(0.002)	(0.002)	(0.002)	(0.004)	(0.004)
Boy2nd	-0.438	-0.438	-0.437	-0.351	-0.373
	(0.002)	(0.002)	(0.002)	(0.004)	(0.004)
Boy1st × *Boy2nd*	0.335	0.335	0.335	0.287	0.299
	(0.003)	(0.003)	(0.003)	(0.005)	(0.005)
Ln[taxable income per		-0.296	-0.198	-0.118	-0.112
capita in district of birth]		(0.004)	(0.004)	(0.006)	(0.005)
Parental education	No	No	Yes	Yes	Yes
R-squared adjusted	0.201	0.205	0.221	0.177	0.204
Model (II)					
Mixed gender	0.100	0.100	0.099	0.062	0.070
	(0.002)	(0.002)	(0.002)	(0.002)	(0.003)
Two girls	0.534	0.534	0.534	0.411	0.440
	(0.002)	(0.002)	(0.002)	(0.003)	(0.004)
Ln[taxable income per		-0.296	-0.198	-0.118	-0.112
capita in district of birth]		(0.004)	(0.004)	(0.006)	(0.005)
Parental education	No	No	Yes	Yes	Yes
R-squared adjusted	0.201	0.205	0.221	0.177	0.204

Note: We assume in Model (II) that the coefficients of *Boy1st* and *Boy2nd* are equal. The reference group in both models is those families with two girls at the first two births. In addition to logarithm of taxable income per capita in district of birth, the set of covariates includes the full set of dummies for urban, parental ages and education, and maternal age at first birth. Standard errors (in parentheses) are heteroscedasticity robust.

Table 3—First-Stage Estimates for Sibsize, Instrumented by Twinning at Second Birth, Linear Models

	First-Born Girls				First-Born Boys			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
				Dependent Variable = $Morethan2$				
$Boy2nd$	-0.221	-0.223	-0.223	-0.222	-0.064	-0.065	-0.064	-0.063
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
$Twin2nd$	0.446	0.331	0.335	0.324	0.578	0.548	0.548	0.537
	(0.005)	(0.006)	(0.006)	(0.006)	(0.004)	(0.006)	(0.006)	(0.006)
$Twin2nd \times Boy2nd$		0.225	0.224	0.223		0.062	0.062	0.063
		(0.008)	(0.008)	(0.008)		(0.008)	(0.008)	(0.008)
Gestation period at second birth (weeks)			0.0020				-0.0007	
			(0.0005)				(0.0005)	
Mean birthweight at second birth (kg)				-0.0108				-0.0152
				(0.0015)				(0.0016)
				Dependent Variable = $Sibsize$				
$Boy2nd$	-0.437	-0.437	-0.436	-0.433	-0.102	-0.102	-0.101	-0.099
	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)
$Twin2nd$	0.645	0.650	0.651	0.629	0.733	0.724	0.717	0.702
	(0.012)	(0.020)	(0.020)	(0.020)	(0.010)	(0.015)	(0.015)	(0.015)
$Twin2nd \times Boy2nd$		-0.009	-0.007	-0.008		0.018	0.020	0.021
		(0.024)	(0.024)	(0.024)		(0.021)	(0.021)	(0.021)
Gestation period at second birth (weeks)			0.0004				-0.0033	
			(0.0009)				(0.0008)	
Mean birthweight at second birth (kg)				-0.0276				-0.0287
				(0.0026)				(0.0023)

Note: Robust standard errors in (.). Additional covariates include parental age, mother's age at second birth, subject's age, birthplace, urban dummy, and logarithm of taxable income per capita in district of birth. Though not reported here, the results change little if an interaction between mean birthweight and *Boy2nd* is included.

Table 4—OLS and 2SLS Estimates

Dependent Variable = HS Completion	First-Born Girls OLS (1)	First-Born Girls 2SLS (2)	First-Born Girls 2SLS (3)	First-Born Girls 2SLS (4)	First-Born Girls 2SLS (5)	First-Born Boys OLS (6)	First-Born Boys 2SLS (7)	First-Born Boys 2SLS (8)	First-Born Boys 2SLS (9)	First-Born Boys 2SLS (10)
	Mediator = $Morethan2$									
$Boy2nd$	-0.0014	-0.0082	-0.0602	-0.0628	-0.0470	-0.0050	-0.0046	-0.0061	-0.0076	-0.0011
	(0.0013)	(0.0040)	(0.0227)	(0.0227)	(0.0229)	(0.0012)	(0.0015)	(0.0093)	(0.0094)	(0.0094)
$Morethan2$	-0.0116	-0.0422	-0.1042	-0.1082	-0.0829	-0.0276	-0.0216	-0.0337	-0.0372	-0.0218
	(0.0015)	(0.0170)	(0.0341)	(0.0341)	(0.0343)	(0.0013)	(0.0134)	(0.0202)	(0.0204)	(0.0206)
$Morethan2 \times Boy2nd$			0.0790	0.0829	0.0599			0.0015	0.0047	-0.0093
			(0.0313)	(0.0313)	(0.0315)			(0.0187)	(0.0188)	(0.0190)
Gestation period at 2nd birth (weeks)				-0.0004					-0.0007	
				(0.0005)					(0.0004)	
Mean birthweight of 2nd birth (kg)					0.0071					0.0061
					(0.0014)					(0.0014)
	Mediator = $Sibsize$									
$Boy2nd$	-0.0044	-0.0116	-0.1031	-0.1098	-0.0784	-0.0057	-0.0050	0.0009	-0.0060	0.0208
	(0.0013)	(0.0053)	(0.0469)	(0.0473)	(0.0479)	(0.0012)	(0.0016)	(0.0368)	(0.0374)	(0.0377)
$Sibsize$	-0.0127	-0.0291	-0.0527	-0.0550	-0.0431	-0.0242	-0.0171	-0.0252	-0.0282	-0.0167
	(0.0009)	(0.0118)	(0.0171)	(0.0173)	(0.0175)	(0.0009)	(0.0105)	(0.0153)	(0.0155)	(0.0157)
$Sibsize \times Boy2nd$			0.0312	0.0334	0.0231			-0.0027	0.0000	-0.0104
			(0.0151)	(0.0153)	(0.0155)			(0.0141)	(0.0143)	(0.0144)
Gestation period at 2nd birth (weeks)				-0.0005					-0.0008	
				(0.0005)					(0.0004)	
Mean birthweight of 2nd birth (kg)					0.0066					0.0058
					(0.0014)					(0.0014)

Note: Same as Table 3. In regressions with "Interact," we interact *Boy2nd* with the mediator, either *Morethan2* or *Sibsize*.

Table 5—Decomposing the Average Total Effect of Having a 2nd-Born Brother on First-Born's High School Completion

	First-Born Girls					First-Born Boys				
	OLS	2SLS				OLS	2SLS			
			Interact	Add GP	Add BW			Interact	Add GP	Add BW
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Average Total Effect	0.0012	0.0013	0.0013	0.0015	0.0005	-0.0032	-0.0032	-0.0032	-0.0032	-0.0036
	(0.0013)	(0.0013)	(0.0013)	(0.0013)	(0.0013)	(0.0012)	(0.0012)	(0.0012)	(0.0012)	(0.0013)
	Mediator = $Morethan2$									
(1) Average Indirect Effect	0.0026	0.0094	0.0232	0.0241	0.0185	0.0018	0.0014	0.0022	0.0024	0.0014
	(0.0003)	(0.0038)	(0.0076)	(0.0076)	(0.0077)	(0.0001)	(0.0009)	(0.0013)	(0.0013)	(0.0013)
(2) Average Direct Effect	-0.0014	-0.0082	-0.0220	-0.0227	-0.0180	-0.0050	-0.0046	-0.0054	-0.0056	-0.0050
	(0.0013)	(0.0040)	(0.0077)	(0.0077)	(0.0077)	(0.0012)	(0.0015)	(0.0018)	(0.0018)	(0.0018)
Controlled Direct Effect	-0.0014	-0.0082	-0.0135	-0.0138	-0.0115	-0.0050	-0.0046	-0.0054	-0.0054	-0.0053
	(0.0013)	(0.0040)	(0.0044)	(0.0044)	(0.0044)	(0.0012)	(0.0015)	(0.0014)	(0.0014)	(0.0014)
	Mediator = $Sibsize$									
(1) Average Indirect Effect	0.0056	0.0128	0.0232	0.0242	0.0190	-0.0032	0.0018	0.0026	0.0029	0.0017
	(0.0004)	(0.0052)	(0.0075)	(0.0076)	(0.0077)	(0.0012)	(0.0011)	(0.0016)	(0.0016)	(0.0016)
(2) Average Direct Effect	-0.0044	-0.0116	-0.0219	-0.0227	-0.0183	-0.0032	-0.0050	-0.0058	-0.0061	-0.0053
	(0.0013)	(0.0053)	(0.0076)	(0.0076)	(0.0077)	(0.0012)	(0.0016)	(0.0020)	(0.0020)	(0.0020)
Controlled Direct Effect	-0.0044	-0.0116	-0.0153	-0.0156	-0.0134	-0.0032	-0.0050	-0.0060	-0.0061	-0.0058
	(0.0013)	(0.0053)	(0.0045)	(0.0045)	(0.0045)	(0.0012)	(0.0016)	(0.0015)	(0.0015)	(0.0015)

Note: Standard errors in (.). The controlled direct effect is evaluated at the mean of the fertility variable. For the list of control variables, see Table 3. "GP" stands for gestation periods. "BW" stands for the mean birth weight of the second birth.

Table 6—Heterogenous Effects for First-Born Girls, by Mothers' Education, Residential Location, and Age at the First Birth

	Hight School Completion (1)	Admitted to University (2)	High School Completion (3)	Admitted to University (4)	High School Completion (5)	Admitted to University (6)
	All First-borns		Mother JHS+		Urban	
Sample mean	0.245	0.176	0.414	0.306	0.296	0.214
Average Total Effect	0.0015	0.0020	0.0001	-0.0002	0.0008	0.0017
	(0.0013)	(0.0012)	(0.0027)	(0.0025)	(0.0023)	(0.0021)
(1) Average Indirect Effect	0.0241	0.0174	0.0366	0.0254	0.0286	0.0316
	(0.0076)	(0.0068)	(0.0112)	(0.0105)	(0.0105)	(0.0096)
(2) Average Direct Effect	-0.0227	-0.0154	-0.0365	-0.0256	-0.0278	-0.0299
	(0.0077)	(0.0069)	(0.0115)	(0.0108)	(0.0107)	(0.0098)
Controlled Direct Effect	-0.0138	-0.0091	-0.0227	-0.0166	-0.0172	-0.0176
	(0.0044)	(0.0039)	(0.0067)	(0.0063)	(0.0063)	(0.0058)
	Mother aged 18-30		Mother aged 31-35		Mother aged 36 or older	
Sample mean	0.241	0.174	0.400	0.293	0.296	0.204
Average Total Effect	0.0017	0.0020	0.0009	0.0065	-0.0668	-0.0448
	(0.0013)	(0.0012)	(0.0100)	(0.0095)	(0.0338)	(0.0305)
(1) Average Indirect Effect	0.0229	0.0154	0.0497	0.0513	-0.0260	-0.0103
	(0.0080)	(0.0072)	(0.0236)	(0.0225)	(0.0381)	(0.0343)
(2) Average Direct Effect	-0.0212	-0.0134	-0.0488	-0.0448	-0.0408	-0.0345
	(0.0080)	(0.0072)	(0.0256)	(0.0243)	(0.0467)	(0.0421)
Controlled Direct Effect	-0.0128	-0.0080	-0.0270	-0.0232	-0.0644	-0.0472
	(0.0046)	(0.0041)	(0.0163)	(0.0155)	(0.0368)	(0.0331)

Note: Estimation is based on mediating variable $Morethan2$. Standard errors in (.). Same as Table 5. JHS indicates junior high school.

Table A1—Mean of first-born singletons who have at least one sibling

	Born in 1978-1984	Born in 1978-1979	
		All	Next sibling born by 1985
	(1)	(2)	(3)
Sample size	821,631	239,107	229,306
Boy-to-girl ratio of 1st born	1.044	1.050	1.051
Boy-to-girl ratio of 2nd born	1.067	1.069	1.068
High school completion rate	0.242	0.194	0.194
College admission rate	0.164	0.134	0.134
More than two children	0.523	0.586	0.601
Complete sibsize	2.685	2.803	2.824
Twins at 2nd birth	0.007	0.006	0.006
Urban (place of birth)	0.339	0.339	0.335
Mother's age at 2nd birth	26.2	25.7	25.5
Mothers' year of birth	1957.3	1955.1	1955.1
Fathers' year of birth	1954.0	1951.7	1951.7
5-year average taxable income per capita in village (thousands)	729.89	727.62	725.22
Mean birthweight of 2nd birth (kg)	3.279	3.289	3.290
Gestation duration of 2nd birth (weeks)	39.62	39.68	39.70
Mothers' highest degree			
College/professional degree or +	0.071	0.064	0.062
High school diploma	0.063	0.053	0.052
Vocational high school	0.191	0.157	0.156
Junior high school	0.261	0.197	0.197
Fathers' highest degree			
College degree or above	0.064	0.062	0.060
Professional degree	0.075	0.067	0.066
High school diploma	0.094	0.088	0.087
Vocational high school	0.183	0.157	0.157
Junior high school	0.233	0.172	0.172

Table A2—Regressions of *Boy2nd* on Family Backgrounds

	Born in 1978-1984	Born in 1978-1979	
Explanatory Variables		All	Next sibling born by 1985
Boy1st	-0.0033	-0.0068	-0.0063
	(0.0011)	(0.0020)	(0.0021)
Urban (place of birth)	0.0014	0.0020	0.0013
	(0.0015)	(0.0028)	(0.0028)
Ln(5-year average taxable income per capita in village (thousands))	-0.0025	-0.0042	-0.0039
	(0.0031)	(0.0059)	(0.0061)
Mother's highest degree			
College/professional degree or +	0.0013	0.0031	0.0048
	(0.0028)	(0.0054)	(0.0056)
High school diploma	0.0007	-0.0005	-0.0002
	(0.0025)	(0.0050)	(0.0052)
Vocational high school	0.0038	0.0084	0.0087
	(0.0017)	(0.0034)	(0.0034)
Junior high school	0.0017	0.0042	0.0049
	(0.0014)	(0.0028)	(0.0029)
Father's highest degree			
College degree of above	0.0031	0.0000	-0.0005
	(0.0029)	(0.0055)	(0.0056)
Professional degree	0.0012	-0.0005	-0.0022
	(0.0025)	(0.0047)	(0.0049)
High school diploma	0.0030	0.0016	0.0000
	(0.0022)	(0.0040)	(0.0041)
Vocational high school	0.0018	0.0049	0.0049
	(0.0017)	(0.0032)	(0.0033)
Junior high school	-0.001	0.0006	-0.0002
	(0.0016)	(0.0030)	(0.0031)
Sample size	821,631	239,107	229,306
R-squared adjusted	0.00000	0.00010	0.00010

Table A3—Regressions of Birth Spacing (Measured in Months) between the First Two Children

	Born in 1978-1984	Born in 1978-1979	
Explanatory Variables		All	Next sibling born by 1985
Girl1st	-0.485	-0.390	-0.306
	(0.069)	(0.120)	(0.074)
Girl1st × *Boy2nd*	0.136	0.104	-0.085
	(0.096)	(0.167)	(0.103)
Boy2nd	0.000	0.014	0.088
	(0.067)	(0.117)	(0.072)
Urban (place of birth)	2.657	1.757	1.115
	(0.066)	(0.117)	(0.072)
Ln(5-year average taxable income per capita in village (thousands))	6.524	5.251	2.868
	(0.145)	(0.261)	(0.157)
Sample size	820,162	238,554	228,753
R-squared adjusted	0.049	0.044	0.063

Table A4—Variable Means for First-Borns, by Twinning at the 2nd Birth

	First-Born Girls		First-Born Boys	
	Twin2nd=0	Twin2nd=1	Twin2nd=0	Twin2nd=1
Sample size	399,078	2,822	417,033	2,698
Outcome variables				
High school completion	0.245	0.246	0.238	0.243
Admitted to university	0.176	0.176	0.153	0.156
More than two children	0.589	1.000	0.453	1.000
Complete family size	2.814	3.399	2.554	3.244
Mean birth weight of 2nd birth (kg)	3.292	2.541	3.276	2.527
Gestation duration of 2nd birth (weeks)	39.64	38.12	39.62	38.10
Urban (place of birth)	0.341	0.358	0.338	0.366
5-year average taxable income per capita in village (thousands)	730.4	736.6	729.3	736.7
Mother's age at 2nd birth	26.2	26.8	26.2	26.8
Mother's year of birth	1957	1957	1957	1957
Father's year of birth	1954	1954	1954	1954
Mother's highest degree				
College/professional degree or +	0.071	0.087	0.071	0.080
High school diploma	0.063	0.065	0.062	0.074
Vocational high school	0.191	0.209	0.191	0.194
Junior high school	0.260	0.263	0.261	0.253
Father's highest degree				
College degree of above	0.064	0.074	0.064	0.071
Professional degree	0.075	0.082	0.075	0.087
High school diploma	0.095	0.100	0.094	0.100
Vocational high school	0.183	0.181	0.184	0.191
Junior high school	0.233	0.237	0.233	0.230

Table A5—First-Stage Estimates for the Interaction between Sibsize and Sibling Gender, Instrumented by the Interaction between Twinning and Sibling Gender, Using Linear Models

	First-Born Girls				First-Born Boys			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	Dependent Variable = $Morethan2$							
$Boy2nd$	0.485	0.482	0.481	0.482	0.425	0.422	0.420	0.421
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
$Twin2nd$	0.285	0.019	0.021	0.011	0.301	0.017	0.017	0.009
	(0.005)	(0.003)	(0.003)	(0.003)	(0.006)	(0.003)	(0.003)	(0.003)
$Twin2nd \times Boy2nd$		0.519	0.520	0.519		0.577	0.579	0.579
		(0.004)	(0.004)	(0.004)		(0.004)	(0.004)	(0.004)
Gestation period at			0.0011				0.0001	
2nd birth (weeks)			(0.0004)				(0.0004)	
Mean birthweight				-0.0119				-0.0098
at 2nd birth (kg)				(0.0012)				(0.0012)
	Dependent Variable = $Sibsize$							
$Boy2nd$	2.607	2.603	2.601	2.603	2.509	2.505	2.503	2.504
	(0.002)	(0.002)	(0.002)	(0.002)	(0.001)	(0.001)	(0.001)	(0.001)
$Twin2nd$	0.323	0.028	0.028	0.011	0.365	0.022	0.021	0.009
	(0.009)	(0.004)	(0.004)	(0.004)	(0.010)	(0.004)	(0.004)	(0.004)
$Twin2nd \times Boy2nd$		0.578	0.581	0.580		0.697	0.697	0.699
		(0.013)	(0.013)	(0.013)		(0.014)	(0.014)	(0.014)
Gestation period at			0.0001				-0.0010	
2nd birth (weeks)			(0.0006)				(0.0005)	
Mean birthweight				-0.0231				-0.0174
at 2nd birth (kg)				(0.0018)				(0.0017)

Note: Same as Table 3.