The Impact of Family Composition on Educational Achievement*

Stacey H. Chen, Yen-Chien Chen, and Jin-Tan Liu

September 23, 2015

Abstract

Intrafamily gender bias, measured by the direct effect (DE) of younger brothers on the first child's education level (given the number of siblings), is understated if parents follow son-preferring fertility stopping rules. Previous literatures have been silent about the relative magnitude of the indirect effect (IE) of younger brothers in a reduced family size because of ill-posed definitional problems. We separately identify the DE and IE of male siblings in an integrated framework. This approach uncovers a new evidence of gender bias in family settings that cannot be derived using conventional methods.

1 Introduction

Gender bias in families has been persisting across generations in many regions. In developing countries such as India, girls get weaned earlier, receive less childcare, and suffer from higher infant mortality.[1] However, other studies find no evidence that females receive less care than males under normal circumstances, even in regions with strongest pro-male bias (Duflo, 2005). For example, seminal work by Deaton (1997, 2003) suggests that there was equal parental spending and vaccination carried out for both genders. Using data from Taiwan, a society with a long tradition of preferring sons over daughters, we also find females are more likely to complete high school or attain university education than males.

There are two possible explanations for this phenomenon: first, son-preferring fertility stopping rules (e.g., Jensen 2005; Barcellos, Carvalho, and Lleras-Muney 2014); and second, females being more enduring than

*S.H. Chen (chens@nber.org): Academia Sinica, Hitotsubashi University, and GRIPS. Y.C. Chen (yenchien@ncnu.edu.tw): Assistant Professor, National Chi-Nan University, Taiwan. Liu (liujt@ntu.edu.tw): 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, John Ham, Yu-Chin Hsu, Amanda Kowalski, Nancy Qian, Francis Vella, Taylor VanderWeele, and participants in several universities and the 2009 NBER Education and Children's Programs. We acknowledge financial support from the Academia Sinica Career Award, the National Science Council (NSC 101-2628-H-001-001-MY3), and the National Health Research Institute.

[1] See Chen, Huq, and D'Souza (1981), Sen (1990), Basu (1989), Ganatra and Hirve (1994), Borooah (2004), Jayachandran and Kuziemko (2011), Barcellos, Carvalho, and Lleras-Muney (2014).

males who were given the same care (e.g., Waldron 1983). The former noted that son-preferring stopping rules not only let males receive more care, but also it consequently discourages having more children. As a result, females appear to receive less care partly because of a larger family size, and not necessarily because of different treatments from parents. On the other hand, the latter explained that because of the more vigorous trait of females, gender differences typically understate the degree of gender bias. Earlier studies on sibling rivalry eliminate the endowment deficits of males by estimating the effect of male siblings on children outcomes (keeping the number of children constant).[2] However, the method of having a constant family size only works for households who do not follow the son-preferring fertility-stopping rule. Therefore, the degree of gender bias is still understated.

Rather than keeping the family size constant, new methods have recently been proposed to address this issue. The most notable work was from Barcellos, Carvalho, and Lleras-Muney (2014) who shutdown the indirect family size channel by restricting their sample to infants under 15 months of age. It was motivated by the fact that mothers are unable to respond to a younger sibling's gender by having more babies. Their result suggested that female infants receive less family resources than males.

Building upon these literatures, we break down the total effect (TE) of having a younger brother on the outcome of older children according to the direct rivalry effect and the indirect family-size effect in an integrated framework. The novelty of this work is that we identify both the DE and IE in one context, and thus, we can compare the magnitude of two effects without restricting to infants or fixing the observed family size.[3] The key idea is that depending on a sibling's sex, potential sibsize can be fixed in counterfactual worlds (where sibling sex is viewed as an assigned experiment); although in reality, observed sibsize cannot.[4] Motivated by VanderWheele's (2013, 2014) causal inference models, we redefine and estimate the DE and IE of sibling sex on educational achievement using Rubin's (1974) counterfactual notation. We address the ill-posed definitional issue in two ways: first, we estimate a selection model with an interaction between randomized sibling sex and endogenous sibsize; and second, we do not evaluate the average DE of having a second-born male on the first child's education at observed sibsize, but at the potential sibsize (as if the next sibling were always male). Moreover, to avoid the ill-posed definitional issue, we avoid muting the IE by requiring an interaction term in the selection model .

This study starts by showing that sibling sex from the second birth is nearly random in Taiwan based on the data from years 1978-1984. In addition, the estimated demand for sons strongly manifests son-preferring fertility stopping rules. We adopt a standard instrumental-variable (IV) method to estimate the selection

[2] See, e.g., Butcher and Case (1994), Kaestner (1997), Garg and Morduch (1998), and Morduch (2000).

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

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

model. The method uses twinning at second birth as an instrument for the fertility choice of having a third child, depending on the initial health conditions of the first child (as in Rosenzweig and Zhang 2009) and a comprehensive list of parental background covariates such as parents' education, residence, age during first birth, and the average income in the neighborhood. With twins, no families are classified as "never-takers" (i.e. if the twins instrument is applied, no family would have only two children) so the twins instrument is not applicable to families with only two children. Nevertheless, the complying families account for about 60% of the population, while the other 40% of the population are "always takers" (i.e., those who would have a third child regardless of the twins instrument switched on or off). Since interpretations of the twins estimates are valid only for the complying families, our results are not conclusive of the general population. If families who insist to have larger family are more traditional and have stronger demand for sons relative to the complying families, then the actual degree of gender bias can be even greater than our estimates.

The evidence of extraordinarily strong demand for sons in Taiwan is seemingly contrary to the near-zero TE on the first child's education when there are younger brothers. By decomposing the TE, our IV estimates suggest that among the complying families, the DE and IE are near zero for first-born males. In contrast, first-born females receive a negative DE and a positive IE, both of which are large and significant. It almost cancels each other out resulting to a near-zero TE. Specifically, it is about 10% of the high school completion or university enrollment rates. In particular, the IE is large and significant for first-born females but has almost no impact on first-born males. This is because having a younger brother considerably reduces the potential sibsize of a first-born female, but has a smaller impact to a first-born male. This phenomenon has been entirely muted by the conventional methods, but it is first uncovered by our empirical strategy.

The remainder of this paper is organized as follows: Section 2 states the concepts and notations used throughout the paper and describes our empirical strategies. Section 3 introduces our data sets, reports descriptive statistics, and examines the exogeneity of a child's sex and the twins instrument. Section 4 summarizes the empirical findings, while Section 5 presents the conclusion of the study.

2 Empirical Strategies

To begin, we describe how DE and IE can be defined and separated.[5] The key is that although potential sibsize can be adjusted in counterfactual worlds, observed sibsize cannot. Under the unconfoundedness assumption, we show that the conventional measure for the direct effect (called "controlled direct effect" CDE) is biased downward. Furthermore, the bias is proportional to the distance between the conditional and unconditional mean of sibsize, given the sibling sex. Our analysis is conditioned implicitly on a set of

[5] See Heckman and Pinto (2015) for literature reviews on causal inference.

covariates, X, which include the first-born child's sex and family background (listed in Table 6).

Where the randomized gender of the second-born sibling D affects the total sibsize M and the first child's outcome Y, let Y_{DM} denote the potential outcome of the first child — given the sibling sex D and sibsize M. We denote M_D as the potential sibsize while Y_D represents the potential outcome given the gender of the next sibling D. The relationship between observed sibsize $M = m$ and potential sibsize $(M_0, M_1) = (m_0, m_1)$ is $m = Dm_1 + (1 - D)m_0$. Specifically, we define the different effects as follow:

- **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** $TE = DE + IE = Y_{1m_1} - Y_{0m_0} = Y_1 - Y_0$.

Notably, the decomposition of the TE is not unique.[6] We focus on the above decomposition since we are interested in the direct rivalry effect on the outcome of females who have younger brothers; and measure the DE by assuming that the potential sibsize m_1 of every family had a male second birth. This interest is similar to the focus of the treatment effect on the treated.

The CDE is constructed by fixing the observed sibsize as $m_0 = m_1 = m$. This approach assumes away the son-preferring fertility stopping rules and implies no IE. Thus, CDE is a conditional total effect restricted only to families with no gender bias in fertility choice.

While the IE captures the impact of younger brothers on the first child entirely through son-preferring stopping rules, the DE captures the other sources by which younger brothers may affect the first child, not through a change in family size. Since a stronger demand for sons induces a larger IE, the IE cannot be omitted for regions where the son-preferring fertility stopping rule is common.

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

$$ADE = \sum_{m \in \mathcal{M}} E[Y_{1m} - Y_{0m} | M = m] Pr\{M_1 = m | D = 1\}, \tag{1}$$

$$AIE = \sum_{m \in \mathcal{M}} E[Y_{0m} | M = m][Pr\{M_1 = m | D = 1\} - Pr\{M_0 = m | D = 0\}],$$

[6] Instead of defining the direct effect as the effect of having a second-born brother holding sibsize fixed at m_1, we could instead fix sibsize at m_0 and define the direct effect as $DE' = Y_{1m_0} - Y_{0m_0}$. Then the indirect effect would fix $D = 1$, instead of $D = 0$, such that $IE' = Y_{1m_1} - Y_{1m_0}$. Both decompositions add up to the same total effect TE.

AIE can be derived by subtracting ADE from ATE since $ADE + AIE = ATE \equiv E[Y_1 - Y_0]$. The identification requires the condition of randomized sibling sex D, which is justified by evidence in Tables 4 and A2. If D is random, then $Pr\{M_d|D = d\}$ can be captured by the mass function of observed sibsize, $Pr\{M|D = d\}$. Our analysis below focuses on the case of binary fertility choice, $M = Morethan2 \in \mathcal{M} = \{0, 1\}$; $M = 1$ if parents have a third child, otherwise M = 0. 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:

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

Intuitively, given a randomized sibling sex, each person's M_0 and M_1 are missing at random. Considering this, we use the conditional expectation of the observed outcome to impute the conditional expectation of the potential outcome.

Because fertility choice is endogenous, we instrument $Morethan2$ by Z, which is the twinning indicator at the second birth. In the incidence of mixed-sex twins, the sex of the second birth D is defined by a Bernoulli random variable with a 50 % probability to be male. To address the endowment deficit of twins, we control the initial health condition of the twins at the second birth that is not affected by the twins sex composition such as the length of gestation periods (see more discussion in Section 3.5). We assume that Z is exogenous given covariates X, as justified by Table 5 (also see Section 3.5). To estimate the target parameters ADE and AIE, we begin with a linear probability model with constant coefficients.

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

The Greek letters are coefficients, and the outcome residual ϵ can be correlated with the selection error u. However, this model is too restrictive because it assumes effect-homogeneity in both observables and unobservables. This model implies $ADE = DE = CDE = \beta_1$, where the CDE is seemingly unbiased but is entirely driven by the functional-form assumption.

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 M by Z and $D \times M$ by $D \times Z$. Second, we replace the outcome error term ϵ with ϵ_{DM} to allow for random effects, whereby individuals self-select depending on idiosyncratic gains from having a brother or having a smaller family. For example, let $Y = \beta_0 + (\beta_1 + \epsilon_1)D + (\beta_2 + \epsilon_2)M + (\beta_3 + \epsilon_3)D \times M + \epsilon$, and $(\epsilon_1, \epsilon_2, \epsilon_3)$ be centered around zero and independent of D and M. Collecting all error terms yields $\epsilon_{DM} = \epsilon + \epsilon_1 D + \epsilon_2 M + \epsilon_3 D \times M$, which is correlated with M but uncorrelated with the random assignment D.[7]

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

Given that D is randomly assigned, ϵ_{0M} and ϵ_{1M} share the same distribution and both are correlated with fertility choice through the selection error u. This model suggests that the ADE, defined in equation (1), can be expressed as

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

The last term is zero because ϵ_{0m} and ϵ_{1m} share the same conditional distribution. If β_1 and β_3 are both identified, then we can identify the ADE; thus, also identifying $AIE = ATE - ADE$:

$$AIE = \beta_2\{E[M|D=1] - E[M|D=0]\}.$$

The ATE in the linear model can be decomposed in an intuitive expression:

$$ATE = \beta_1 + \beta_3 E[M|D=1] + \beta_2\{E[M|D=1] - E[M|D=0]\}.$$

Notably, given the sibling sex, both effects are affected by the conditional probability of fertility choice. If younger brothers discourage fertility, then $E[M|D=1] - E[M|D=0] < 0$. On the other hand, if β_2 is also negative, then $AIE > 0$ offsets a negative ADE.

By a less restricted model (2), the conventional measure for the ADE is

$$CDE = \beta_1 + \beta_3 E[M]. \tag{4}$$

This is a biased measure for ADE, unless fertility choice is independent of the sex composition of the previous children. If parents adopt son-preferring fertility stopping rules, then $E[M|D=1] < E[M]$ and the bias of

[7] In our IV estimation, the interaction between D and M allows for observed heterogeneity in the effects of family size by gender of the first born, and it needs to be instrumented too, using another first-stage equation: $D \times M = \gamma_0 + \gamma_1 D + \gamma_2 Z + \gamma_3 D \times Z + v$.

CDE increases with the difference between $E[M|D = 1]$ and $E[M]$. We note that the conventional method is entirely silent about the identification of the AIE.

We estimate β_1 and β_3 by standard IV methods using the twins instrument $Z = 0$ or 1. The IV estimates provide a causal interpretation only for complying families, whose sibsize would rise with twinning at the second birth ($M(Z = 1) > M(Z = 0)$). Thus, estimation using all families (including those who would have a third child even in the absence of twins), may give different results from the IV estimates. Brinch, Mogstad, and Wiswall (2015) suggest that local average treatment effect (LATE) of family size using the twins instrument differ substantially from the population average effects. Similarly, the local average DE and IE likely differ from ADE and AIE for the general population. Precisely, the LATE analogs for ADE and AIE are

$$LADE = \beta_1 + \beta_3 E[M|D = 1, M(1) > M(0)],$$
$$LAIE = \beta_2\{E[M|D = 1, M(1) > M(0)] - E[M|D = 0, M(1) > M(0)]\}.$$

Unless the compliance rate is nearly 100%, the LATE results cannot identify the ADE and AIE for the general population. The first-stage estimates in Table A3 imply the compliance rate, given a second-born son, to be about 56 to 62%.[8] Because there are no "never-takers" with the twins instrument, all of the non-complying families (about 38-44%) are "always-takers" who would have a third child anyway, regardless of having twins or a singleton at second birth. Our estimated local ADE and AIE represent around 60% of the general population.

3 Data and Descriptive Analysis

Identifying the 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 of completed families and children's educational attainment up to the late teens. To fulfill this requirement, we link two Taiwanese national administrative datasets, Birth Registry and University Entrance Test records.

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. The data also contains everyone's identifier to allow us to link all children to mothers. We restrict our data to 929,754 mothers whose first birth was at the age of 18 or older (prior to January 4, 1985, when the Eugenics Protection Law began to be enforced). Although prenatal sex testing by ultrasound

[8] We derive the compliance rate by $Pr\{M(1) > M(0)|D = 1\} = E[M|Z = 1, D = 1] - E[M|Z = 0, D = 1]$

was introduced in Taiwan 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 order (Lin, Liu and Qian 2014).

Birth Registry has detailed categorical information about parental education. Because the years of education in general tracks versus vocational tracks are not comparable, we capture the variation in parental education by using five indicators: university degree or higher, professional training degree, high school diploma, vocational high school diploma, and junior high school diploma. The excluded category, primary school or lower, is the reference group. We further include family socioeconomic status by merging the data with per capita taxable income by the district of birth.

To measure the sibling sex composition of completed families, we trace all births of 929,754 mothers for 15 to 22 years until 1999. No mother in our data had a child in either 1998 or 1999, so the measures of completed family size and sibling sex composition are accurate. Taiwan has no birth-control policy promulgated, so our data are not distorted by under-reported female births induced by forceful birth-control policies.

Table 1 summarizes the distribution of the 929,754 families by number of children. To causally link child education to sibling sex composition (depending on the birth order), we focus on the education of 821,631 first-born singletons from 2+ families (which account for 88% of all families). For families with two or more children, the sex ratio of males to females drop rapidly with the number of children. The sex ratio goes from 1.4 for families with two kids to one or less for those with three or more kids. This is consistent with the notion that parents stop having children when they have a son. It also suggests that child gender and family size are both endogenous. However, we argue below that child gender is close to a random assignment after controlling for a list of comprehensive covariates, including parental education and location of residence.

3.1 A First-Born Son Reduces Family Size

After having the first child, the decision whether to have more children depends on the sex of that child. On average, first-born males have 0.27 fewer siblings than first-born females, irrespective of including demographic and socioeconomic covariates. This is about 10% of the average sibsize of all families (2.7). We report these results in the top panel of Table 2. The effect of having a first-born son on sibsize is greater among families in rural areas, or when the mother's education level is less than junior high school. The estimated effect increases to about 0.30 with standard errors as small as 0.002.

Because a first-born son significantly reduces the chances of having a second child, the families with a first-born son or first-born daughter who has a certain number of siblings are not comparable because the former group probably has preferences for larger families. Thus, our analysis separates 419,731 first-born sons from 401,900 first-born daughters out of the 821,631 families with two or more children.

3.2 Education and Characteristics of First-Born Children

We acquired education data from the University Entrance Test records of 1996 to 2003 when the first-born just turned 18. The data include two sets of tests: general tests (conducted in February during the high school senior year) and union entrance tests (conducted in July after high school graduation). These tests offer two distinct channels for university education: first, students can apply for university admissions using their general test scores and skip the tests in July. If their application results are unsatisfactory, students can forgo early admissions and take the union entrance tests in July after graduation. The indicator for university admission in our study is based on both channels. 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. Since 1928, when the first university was founded, the brightest students in Taiwan have attended public universities. During our sample years from 1996 to 2003, tuition and fees in public universities were about 14% of the yearly family income, whereas the cost of attending private colleges was about 25%.

First-born children are classified as receiving an intervention if the next sibling is a brother rather than a sister ($Boy2nd = 1$). In the control group, the next sibling is a sister ($Boy2nd = 0$). If the second birth results in mixed-sex twins or triplets, then we randomize sibling sex by the fraction of males from the birth. We assign 1 to sibling sex to those with the probability of a male in mixed-sex twins (0.5), the probability of a male in triplets with one male (0.33), and the probability of a male in triplets with two males (0.66); otherwise, we assign 0. Overall, there are 424,166 first-born children in the treatment group who have a second-born brother, while there are 397,465 in the control group who have a second-born sister.

Table 3 reports the statistics of first-born outcomes and characteristics by sibling sex composition. Statistics show that the covariates of the treatment and control groups ($Boy2nd = 1$ and $Boy2nd = 0$) is balanced, with nearly identical family backgrounds and socioeconomic status. About a quarter of the first-born children completed high school, and only 15-18% enrolled in universities.

First-born females 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 females have 0.54 more children than those with two males, and 0.43–0.44 more children than those with a mixed-sex composition.

Unlike American parents who prefer a mixed-sex composition, Taiwanese parents strongly favor multiple sons, as confirmed in the next subsection. As Table 3 indicates, Taiwanese families with two females are 28 percentage points more likely to have a third child than those with two males. This is extraordinarily large, compared to the same differential in the United States and Israel, which is less than two percentage points (Ben-Porath and Welch 1976, Angrist and Evans 1998, Angrist, Lavy and Schlosser 2010).

3.3 Demand for Multiple Sons — Son-Preferring Fertility-Stopping Rules

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 acted 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 old-age social security (not based on employment) in Taiwan began in 2008, 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, such as the *Analects* (ca. 479 BCE), systematically justifies the demand for multiple sons. The *Analects* were at the core of the educational curriculum in Imperial China for more than two millennia.[9] Until now, Confucianism still remains as a dominant component of the educational curriculum in Taiwan.

We report the demand for multiple sons in the middle and bottom panels of Table 2, where we estimate the effect of a change in sibling sex composition on sibsize in the 2+ families, given observed family backgrounds. 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 middle panel shows that having a son, regardless of the birth order, decreases sibsize by 0.43-0.44 person. Because birth order is not important in explaining the demand for sons in our data, we further 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 with families with two males, those with two females have about 0.53 more

[9] 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.

children, and those with mixed-sex composition have about 0.1 more. These estimates are extraordinarily large, since they account for approximately 20% and four percent, respectively, of the average sibsize (about 2.7). If the child was born in a rural area or the mother has no junior high school diploma, then the level of these estimates further increases by about 11% to 18%. These results are robust and precise, whether or not we include parental education, per capita taxable income in the district of residence, or both.

3.4 Testing for Exogeneity of Sibling Sex Composition

Although the presence of sex-selective abortion is neither observable nor testable in the data, we examine the exogeneity of sibling sex composition of the 2+ families in four ways. First, 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 (between 1.05-1.08) that demographers consider normal on the basis of historical evidence (Johansson and Nygren 1991).

Second, we compare demographics of the full sample (born between 1978-1984) 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 socioeconomic status, 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, we regress the sex of the second child with the sex of the first child after including a list of observed family background variables. We find in Table A2 that the R-square adjusted is close to zero. The implied F statistic is below the critical value at the 99% significance level. Having a first-born female is associated with a 0.33% significant increase in the ratio of males to females at the second birth. 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 4 shows that the estimated coefficient of the interaction between $Girl1st$ and $Boy2nd$ is only four days, which is statistically insignificant. Therefore, we reject the hypothesis that after having given birth to a female first, the mother tends to spend more time trying to bear a male relative to a female. Our statistical results suggest that sex-selective abortion is not a concern among children from the first two births in our data.

3.5 Exogeneity of the Twins Instrument

Exogeneity of the twins 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 DE 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. Additionally, if parents preferring singletons are in favor of sons, the diversion of family resources might be greatest from female 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 males.

Although exogeneity of the twins instrument is not testable, we examine whether the occurrence of twins can be explained by family backgrounds such as parental education, place of birth, or the average taxable income in the district of birth. In Table A5, we first compare the outcomes and family backgrounds of first-born children between families with a twin versus those without a twin pair during the second birth. While twinning at the second birth increases family size sharply by at least 0.5 children, it appears to have no effect on first-born females' education. We note that twinning is associated with a two-percent increase in first-born males' high school completion and university attainment, but this is likely due to an older or more educated mother. On the other hand, it is known in medical literature that women are naturally more likely to conceive twins as a result of greater fluctuations in hormone levels. Hence, more educated women are more likely to take fertility inducing drugs that also increase the incidence of twinning.

We show in Table 5 that after controlling for maternal age during the second birth, neither parental education nor socioeconomic status can explain the incidence of twinning. As the incidence of twining rises with maternal age, it also rises over time. Column (1) suggests that the birth year of the second birth can explain only up to 0.015% of the variation in the incident of twinning. In Columns (3) and (4), we include the parents' education, urban residence, and socioeconomic status (captured by log taxable income per person in the neighborhood). The indicators for urban residence and parental education levels are marginally significant, but only increase the adjusted R-square by 0.00002. We further add in Column (4) the full set of dummies for maternal age at the second birth, which explains the additional 0.007 percentage points of variation in the incidence of twinning. This also considerably reduces the explanatory power of parental education and socioeconomic variables. Now, neither socioeconomic variables nor parental education levels are important explanatory variables. Overall, the F statistic cannot reject the hypothesis that the coefficients for these background variables are jointly zero. These results suggest that the birth of twins is not related directly to parents' education or socioeconomic status.

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 by fixing the birth weight of the second birth in addition to family backgrounds, the only channel through which twinning at the second birth can affect the first-born child's education is through 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.

Because boys are heavier than girls at birth on average, part of the sibling-sex effect may be mistaken for a birth-weight effect if we include the mean birth weight of the second birth as a control variable. Our results in Table III. C in Chen, Chen, and Liu (2014) suggest that adding the mean birth weight of the second birth leads to a 20% decrease in the 2SLS estimates of the sibling-sex effects among first-born females. The downward adjustments for first-born males are smaller and imprecise. The 2SLS estimated family-size effects are also adjusted downward for first-born females by 20% or higher. 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-sex effects. Thus, inclusion of the mean birth weight may open up another causal channel — from sibling sex, 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 sex may indirectly affect the first-born outcomes through changes in sibling birth weight, in addition to changes in sibsize.[10]

One alternative control for the initial health of the second-born is the length of gestation periods, which is not affected by gender. As statistics in Table 3 show, the gender gap in gestation duration is only 1% of the average duration and it is statistically insignificant. The result in Column (3) of Table A3 suggests that an additional one week in the second-born gestation period increases the likelihood of having a third child by 0.002 (SE=0.0005) if the first born is female. The second-stage result in Tables 6 indicates that

[10]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) and Currie (2009) for reviews of the literature. To formally estimate the indirect effect via changes in sibling birth weight, we need to instrument birth weight.

the second-born gestation has almost no impact on the first child's education. The estimates based on gestation duration cannot support the conjecture that parents are in favor of singletons over twins due to their difference in initial health.

4 Results

Using 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 for having a third child, second-stage estimates of the human capital formation function of the first-born, and decomposition of the TE into DE and IE. We show that conventional methods (using either the TE or the CDE) might have systematically understated the degree of gender bias, primarily due to the muted interaction between sibling sex composition and the decision to have more children.

It is important to emphasize that even if family size were exogenous, omitting the interaction term would still cause ill-posed definitional problems for both DE and IE. This definitional issue should be treated first, before tackling the problem of endogeneity in family size. Additionally, we address the issue of health deficits of twins by conditioning on an initial health condition that is not related to gender.

4.1 First-Stage Estimates

We instrument the decision to have a third child, using the occurrence of twins at the second birth. The first-stage results are very strong for both of the fertility choice variables ($Morethan2$ and $Sibsize$) and for their interaction with sibling sex $Boy2nd$. The estimates in the top panel of Table A3 suggests that a twin birth increases the probability of having more than two children by 33-61 percentage points (see columns (2) and (6), where the 61 percentage points increase is derived from 0.548+0.062 for families with $Boy2nd = 1$). Those in the bottom panel show that a twin birth increases the completed sibsize by more than 0.6 children. These estimates are robust and significant, irrespective of including initial health of the second birth.

The first-stage estimate decreases with the number of sons because of strong demand for (multiple) sons. If the first two births are both males, 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 male. The number goes further down to 34 percentage points if they are all females because some parents who have no sons keep trying, whether or not they give birth to twins. Compared to families with all females, a family with twins is more likely to push parents with at least one male 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 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 A3 indicate that having a second-born male increases the effect of twins on sibsize only by 0.02 or less, which 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 A4 shows). The coefficient of the interaction is about 0.5-0.7, with very small standard deviation. These figures are robust and significant across various models, suggesting that we do not have to be concerned about weak instruments.

4.2 OLS and 2SLS Results

In Table 6, we regress the first-born's completion of high school on family composition variables (including sibsize and sibling sex). OLS methods considerably understate the family size effects on the first-born females' education because of the omitted-variable bias and the omitted interaction term. The downward bias is much larger among first-born females than first-born males. OLS-estimated coefficients of $Morethan2$ and $Sibsize$ for first-born females are only about -1 percentage point. However, unobserved bias against first-born daughters can be greater in large families and in the presence of a younger brother (that is, the outcome residue is correlated with sibsize and the omitted interaction between sibsize and $Boy2nd$), so the family-size effect on first-born females' education may be understated.

We instrument family size using the occurrence of twins at the second birth. In order to show how the 2SLS results change when adding the interaction between sibsize and sibling sex, holding the set of instruments (and compliers) fixed, we include interactions between the twins instrument and sibling sex as an additional instrument for family size also in the model without interactions between sibsize and sibling sex.

Compared with the OLS results, the 2SLS estimated coefficient of family size is considerably increased in magnitude for first-born females as expected, and little is changed for first-born males. These results can be found in Columns (2) and (6). As $Boy2nd$ is not endogenous in our data (recall Section 3.4), a large change in the coefficient of $Boy2nd$ in Column (2) after instrumenting fertility choice is noteworthy. The 2SLS estimated coefficient of sibsize for first-born females also rises substantially. It is likely that for first-born females, the sibling-sex 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 sex to interact in the regression analysis.

As Columns (3) and (7) show, adding an interaction term between family size and $Boy2nd$ considerably changes the 2SLS result, particularly for first-born females. The coefficient of the interaction term is large and significant, and the coefficient of $Boy2nd$ rises at least tenfold. In contrast, both coefficients for first-born males are smaller and insignificant.

The coefficient of *Boy2nd* on child outcomes generally cannot be interpreted as the DE of sibling sex on education (as we have emphasized in Section 2). In contrast, the 2SLS estimated coefficient of fertility choice still has important causal interpretations for the family-size effect 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 of the firstborn falls by about 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 more than 40% and 20% of the high school completion rate, respectively. The largest family size effect appears among first-born females whose next sibling is also a female. It is because parents who would keep on trying after having all females in the first two births are most likely to invest only in the later-born son, compared with those who stop.

By contrast, if there are one or more sons in the first two births, then the family-size effect is reduced and becomes imprecise. As Columns (3)(4) and (7)(8) suggest, having more than two children with at least one male decreases the high school completion rate by about 3 percentage points (with standard errors being around 0.02-0.03). Although these estimates are not small, they are too imprecise to be conclusive.

4.3 Main Results: Decomposition

With the extremely strong demand for sons, it is perhaps surprising that on average, Taiwanese females are more likely than males to complete high school and enroll in the university. Also, the ATE of having a second-born son on whether the first-born female completes high school is positive or nearly zero (as the first row of Table 7 shows). In contrast, the same effect is negative for first-born sons. These statistics might be seen as evidence for the absence of rivalry effects of male siblings on Taiwanese females, even with exceedingly strong demand for sons. The key to explaining this puzzle is the presence of positive AIEs owing to son-preferring fertility stopping rules. Reduced sibsize, after having a subsequent brother, allows more family resources to be invested in the first-born females' education. Since the AIEs run in the opposite direction of the ADEs, the ATEs are close to zero. We expand on these results below.

Columns (1) to (4) of Table 7 report our decomposition results for first-born females. The estimated ADEs and AIEs rise considerably from Column (1) to Column (2), after we address the endogeneity of fertility choice. The adjustments go further after we add in Column (3) an interaction between sibsize and sibling sex. This indicates a great deal of heterogeneity in the sibling-sex effects across various sibsizes. The interaction term should not be omitted from the model.

Unlike the large adjustments for endogenous sibsize among first-born females, these adjustments among first-born males are small and insignificant (as Columns (5) and (6) show). This is because first-born males have considerably smaller families than first-born females, regardless of whether their parents opt for child quality over quantity. 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 ADEs and AIEs of sibling sex are both nearly zero, interacting sibsize with sibling sex has almost no impact on the estimated results.

We note that the gap in estimates between the CDE and the ADE is nearly zero for first-born males, while it is much wider for first-born females. After we address endogenous sibsize in Column (2) and add the interaction term in Column (3), the 2SLS estimated CDE is much smaller than the ADE for first-born females, while it is about the same for first-born males. This contrast is due to the fact that the CDE is evaluated at the unconditional average sibsize (as in equation (4), under the assumption that sibsize does not change with sibling sex), while the ADE is evaluated at the average sibsize depending on the second child being a male (see equation (3)). Owing to son-preferring fertility stopping rules, having a second-born son as opposed to a second-born daughter reduces the likelihood that parents of first-born females will keep trying. As a result, the average sibsize of first-born females, conditional on having a second-born brother, is smaller than the unconditional mean. In contrast, for first-born males, the conditional and unconditional means of sibsize are almost equal, so the bias of the CDE is close to zero.

Overall, sibling-sex effects are much smaller on first-born males than on first-born females. The AIE of sibling sex on first-born females' education is more than 10 times that of first-born males' education. On the other hand, the ADE is more than four times that of the first-born males' education. This evidence points to a very strong pro-male bias, much stronger than what the CDE has indicated, and the opposite of what the ATE has suggested.

The 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 (see Tables 7 and 8). Indeed, the choice of a mediator should not alter the magnitude of the ADE of a given intervention on outcomes.

Decomposition results for another important education outcome — university admission at age 18 — show similar patterns, as Table 8 presents. The ATE of having a second-born brother on the first child's university enrollment is positive or close to zero for both genders. On the basis of the estimated ATE, gender bias seems to be absent or at least not against females. Only after we divide the ATE into ADE and AIE does gender bias become evident. Unlike first-born males, whose ADE and AIE are both close to zero, first-born females receive a boost of 1.7 percentage points in their AIE and suffer a loss of 1.5 percentage points in their ADE. Both estimates are statistically significant. These results are considerable in magnitude, since they account for 8-10% of first-born females' university enrollment rate, which is about the same proportion as the effects on first-born females' rate of high school completion reported earlier.

4.4 Heterogeneous Effects of Sibling Sex Composition on Education

Gender bias against first-born females appears to be most evident in attaining university education if they were born in urban areas. In contrast, we find little evidence of gender bias against first-born males' education. If any, it is limited to those born to less-educated mothers residing in rural areas (the magnitude being no more than 5% of the sample mean). These results are suggested in Table 9 by the estimated ADE of having a second-born brother on first-born education, divided by the group-specific average of the educational attainment.

Columns (1) to (4) show that the ADE of having a son on first-born females' university attainment is strongest if they were born to mothers residing in urban areas (the estimated ADE is about -14% of the university enrollment rate). In contrast, the same effect in rural areas is less than 3% of the enrollment rate. On first-born females' high-school education, the ADE has less variation, ranging between 7-9% of the completion rate, regardless of place of birth or maternal education levels. All of these negative ADE are canceled out by a strong positive AIE. As a result, the ATE of having a brother in each group is either positive or close to zero. First-born males are not affected as strongly as first-born females by sibling sex, neither directly nor indirectly, as Columns (5) to (8) show. Both effects are statistically insignificant or close to zero with small standard errors. Contrary to first-born females always having ADEs being entirely canceled out by AIEs, first-born males have ADEs exceeding AIEs on their high school completion. Consequently, the ATE of having a younger brother is positive or nearly zero on the education of first-born females, but negative for education of first-born males. Without decomposition, the ATE gives a wrong impression that gender bias is absent in Taiwan.

5 Conclusion

Mixed evidence of intrafamily gender bias in regions with strongest son preferences is likely caused by son-preferring fertility stopping rules adopted in many families. The ADE of younger brothers on older children can be offset by the indirect benefits from reduced family size. Although this is clear conceptually, little empirical evidence uncovers the relative magnitudes of the ADEs and AIEs. In this study, we clarify the ill-posed definitional issue in measuring the ADEs and AIEs using an integrated framework. We decompose the ATE to ADE and AIE in one context so we can compare their relative importance.

Our key result is that both DE and IE account for one-eighth to one-tenth of the average educational achievement of first-born females, in opposite directions. This leads to a near-zero ATE. Additionally, the effect of one additional sibling lowers her opportunity for a university education by 10 percentage points (about two-fifth of the average university admission rate). In contrast, neither the number of siblings nor the gender of the next sibling has a noticeable effect on a first-born male, 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. As in any study, this work has some limitations. First, our empirical setting is restricted to families who have at least two children. Second, our identification strategy does not allow us to identify the DE and IE for families who would have a third child anyway, regardless of the presence of second-born twins.

Nevertheless, the estimated magnitudes of intrafamily gender bias is important for policy. Unlike the previous evidence of gender bias mostly focusing on infant females, our results show that intrafamily gender bias has a sizable negative impact on first-born females' high school or university education, although it is mostly indirectly offset by parents' fertility stopping rules. As China recently started to relax the one-child policy, our results suggest that even if intrafamily gender bias against females is likely to continue, some will be offset by the indirect benefits created by allowing couples to freely choose their desirable number of children. Our findings have important implications for model specification strategies for separating causal channels, definitional issues of intrafamily gender biases, and long-term impacts of intrafamily gender bias on female adolescents.

References

Almond, Douglas, Kenneth Y. Chay, and Davis S. Lee. 2005. "The Costs of Low Birth Weight," *Quarterly Journal of Economics* 120(3): 1031-1083.

Ananat, Elizabeth O. and Guy Michaels. 2008. "The Effect of Marital Breakup on the Income Distribution of Women with Children?" *Journal of Human Resources,* 43(3): 611-629.

Angrist, Joshua D. and W. Evans. 1998."Children and Their Parents' Labor Supply: Evidence from Exogenous Variations in Family Size." *American Economic Review* 88(3): 450-477.

Angrist, Joshua D., Victor Lavy, and Analia Schlosser. 2010. "Multiple Experiments for the Causal Link between the Quantity and Quality of Children," *Journal of Labor Economics* 28(4): 773-823.

Barcellos, Silvia Helena, Leandro S. Carvalho, and Adriana Lleras-Muney. 2014. "Child Gender and Parental Investments In India: Are Boys and Girls Treated Differently?" *American Economicy Journal: Applied Economics* 6(1): 157-189.

Ben-Porath, Yoram and Finis Welch. 1976."Do Sex Preferences Really Matter?" *Quarterly Journal of Economics* 90(2): 285-307.

Brinch, Mogstad, and Wiswall, 2015. "Beyond LATE with a Discrete Instrument: Heterogeneity in the Quantity-Quality Interaction of Children." Mimeo.

Black, Sandra E., Paul J. Devereux and Kjell G. Salvanes. 2005. "The More the Merrier? The Effect of Family Size and Birth Order on Children's Education." *Quarterly Journal of Economics* 120(2): 669-700.

Butcher, F. Kristin and Anne Case. 1994."The Effect of Sibling Sex Composition on Women's Education and Earnings." *Quarterly Journal of Economics* 109(3): 531-563.

Chen, Stacey H., Yen-Chien Chen, and Jin-Tan Liu. 2014. "The Impact of Family Composition on Educational Achievement." NBER Working Paper No. 20443.

Croll, Elisabeth. 2000. *Endangered Daughters: Discrimination and Development in Asia,* London and New York: Routledge.

Currie, Janet. 2009. "Healthy, Wealthy, and Wise: Socioeconomic Status, Poor Health in Childhood, and Human Capital Development." *Journal of Economic Literature* 47(1): 87-122.

Dahl, Gordon and Enrico Moretti. 2008. "The Demand for Sons: Evidence from Divorce, Fertility and Shotgun Marriage." *Review of Economic Studies* 75(4): 1085-1120.

Das, Narayan. 1987. "Six Preference and Fertility Behavior. A Study of Recent Indian Data." *Demography* 24: 517-530.

Deuchler, Martina. 1992. *The Confucian Transformation of Korea: A Study of Society and Ideology.* Cambridge MA: Council on East Asia Studies, Harvard University (Harvard-Yenching Institute Monograph Series 36).

Duflo, Esther. 2005. "Gender Equality in Development." Bureau for Research in Economic Analysis of Development (BREAD) Policy Paper 011.

Frolich, Markus and Martin Huber. 2014. "Direct and Indirect Effects - Casual Chains and Mediation Analysis with Instrumental Variables." IZA #8280.

Garg, Ashish and Jonathan Morduch. 1998. "Sibling Rivalry and the Gender Gap: Evidence from Child Health Outcomes in Ghana." *Journal of Population Economics.* 11(4): 471-493.

Heckman, James J. and Rodrigo Pinto. 2015. "Econometric Mediation Analyses: Identifying the Sources of Treatment Effects from Experimentally Estimated Production Technologies with Unmeasured and Mismeasured Inputs," *Econometric Reviews*34(1-2): 6-31.

Huber, Martin. 2014. "Identifying Causal Mechanisms (Primarily) Based on Inverse Probability Weighting." *Journal of Applied Econometrics* 29(6): 920-943.

Jayachandran, Seema and Ilyana Kuziemko. 2011. "Why do Mothers Breastfeed Girls Less Than Boys? Evidence and Implications for Child Health in India." *Quarterly Journal of Economics* 126(3): 1485-1538.

Jensen, Robert. 2003 "Equal Treatment, Unequal Outcomes? Generating Sex Inequality through Fertility Behavior." Harvard University.

Johansson, S. and O. Nygren. 1991. "The Missing Girls of China - A New Demographic Account," *Population and Development Review* 17(1): 35-51.

Kaestner, Robert. 1997. "Are Brothers Really Better? Sibling Sex Composition and Educational Achievement Revisited." *Journal of Human Resources*, 32(2): 250-284.

Lin, Ming-Jen, Jin-Tan Liu and Nancy Qian. (2014), "More Missing Women, Fewer Dying Girls: The Impact of Sex Selective Abortion on Sex at Birth and Relative Female Mortality in Taiwan." *Journal of the European Economic Association,* 12(4): 899-926.

Lundberg, S. J. and E. Rose. 2002. "The Effect of Sons and Daughters on Men's Labor Supply and Wages." *Review of Economics and Statistics*, 84(2): 251-268.

Morduch, Jonathan. 2000. "Sibling Rivalry in Africa." *American Economic Review*, 90(2): 405-409.

Mogstad, Magne and Matthew Wiswall. 2014. "Testing the Quantity-Quality Model of Fertility: Estimation Using Unrestricted Family Size Models." *Quantitative Economics, forthcoming*

Parish, William and Robert Willis. 1993. "Daughters, Education and Family Budgets: Taiwan Experiences." *Journal of Human Resources*, 28(4): 863-898.

Pearl, Judea. 2001. "Direct and Indirect Effects." *Proceedings of the Seventeenth Conference on Uncertainty in Artificial Intelligence*, Technical Report R-273-UAI.

Rose, Elaina. 2000. "Gender Bias, Credit Constraints and Time Allocation in Rural India." *Economic Journal,* 110(465): 738-58.

Rosenzweig, Mark R. and Kenneth I. Wolpin. 1980 "Testing the Quantity-Quality Fertility Model: The Use of Twins as a Natural Experiment." *Econometrica*, 48(1): 227-240.

Rosenzweig, Mark R. and Junsen Zhang. 2009. "Do Population Control Policies Induce More Human Capital Investment? Twins, Birth weight and China's 'One-Child' Policy," *Review of Economic Studies*, 76(3): 1149-1174.

Rubin, Donald B. 1974. "Estimating Causal Effects of Treatments in Randomized and Nonrandomized Studies," *Journal of Educational Psychology*, 66(5): 688-701.

VanderWeele, Tyler J. 2013. "A Three-Way Decomposition of a Total Effect into Direct, Indirect and Interactive Effects," *Epidemiology*, 24(2): 224-232.

VanderWeele, Tyler J. 2014. "A Unification of Mediation and Interaction: A 4-Way Decomposition," *Epidemiology*, 25(5): 749-761.

Wei, Shang-Jin and Xiaobo Zhang. 2011. "The Competitive Saving Motive: Evidence from Rising Sex Ratios and Savings Rates in China," *Journal of Political Economy* 119(3): 511-564.

Table 1: Family Characteristics by Completed Number of Children

	All	Number of Children									
	Families	1	2+	2	3	4	5	6	7	8	9+
Frequency	929,754	108,123	821,631	392,244	322,914	85,271	16,584	3,536	795	202	85
Percentage	100	11.6	88.4	42.2	34.7	9.17	1.78	0.38	0.09	0.02	0.01
Sex ratio (boys/girls)	1.07	1.30	1.04	1.39	0.94	0.51	0.39	0.43	0.45	0.59	0.55
Urban (place of birth)	0.36	0.50	0.34	0.41	0.29	0.23	0.20	0.19	0.21	0.18	0.31
Taxable income per capita in district (thousands)	739	805	730	765	705	678	664	657	654	634	679
Mother's year of birth	1957	1955	1957	1957	1958	1958	1958	1959	1959	1959	1959
Father's year of birth	1954	1951	1954	1953	1955	1955	1955	1954	1954	1953	1954
Mother's highest degree											
College/professional degree+	0.08	0.14	0.07	0.11	0.04	0.02	0.01	0.01	0.01	0.00	0.00
High school diploma	0.07	0.09	0.06	0.08	0.05	0.03	0.02	0.02	0.02	0.01	0.00
Vocational high school	0.19	0.21	0.19	0.24	0.16	0.11	0.08	0.06	0.06	0.04	0.05
Junior high school	0.25	0.21	0.26	0.24	0.28	0.27	0.25	0.24	0.21	0.16	0.20
Father's highest degree											
College degree+	0.07	0.13	0.06	0.10	0.03	0.01	0.01	0.01	0.01	0.00	0.00
Professional degree	0.08	0.10	0.07	0.10	0.06	0.03	0.02	0.02	0.02	0.00	0.00
High school diploma	0.10	0.12	0.09	0.12	0.08	0.06	0.05	0.05	0.04	0.04	0.06
Vocational high school	0.18	0.17	0.18	0.20	0.18	0.14	0.12	0.11	0.08	0.06	0.06
Junior high school	0.23	0.18	0.23	0.20	0.26	0.27	0.26	0.24	0.22	0.21	0.19

Note: We exclude families having multiple babies at the first birth. We exclude families with the father younger than 18, the mother younger than 18 at the first birth, or missing information about child birth year or per capita taxable income in district.

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

Depedent Variables =*Sibsize*		Add Income	Add Parents' Edu.	Born in Rural	Mother Less than JHS
			All Families		
Boy1st	-0.266	-0.266	-0.266	-0.298	-0.291
	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)
Ln(taxable income per		-0.309	-0.309	-0.673	-0.441
capita in district of birth)		(0.005)	(0.005)	(0.006)	(0.008)
R-squared adjusted	0.16	0.16	0.16	0.06	0.13
Sample size	929,754	929,754	929,754	615,080	596,822
		Families with Two or More Children			
Model (I)					
Boy1st	-0.432	-0.432	-0.432	-0.473	-0.466
	(0.002)	(0.002)	(0.002)	(0.003)	(0.003)
Boy2nd	-0.438	-0.438	-0.437	-0.479	-0.471
	(0.002)	(0.002)	(0.002)	(0.003)	(0.003)
Boy1st × *Boy2nd*	0.335	0.335	0.335	0.359	0.354
	(0.003)	(0.003)	(0.003)	0.0041	0.0040
Ln(taxable income per		-0.296	-0.198	-0.264	-0.317
capita in district of birth)		(0.004)	(0.004)	(0.006)	(0.007)
R-squared adjusted	0.20	0.21	0.22	0.18	0.21
Model (II)					
Mixed gender	0.100	0.100	0.099	0.117	0.114
	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)
Two girls	0.534	0.534	0.534	0.593	0.583
	(0.002)	(0.002)	(0.002)	(0.003)	(0.003)
Ln(taxable income per		-0.296	-0.198	-0.264	-0.317
capita in district of birth)		(0.004)	(0.004)	0.0059	0.0070
R-squared adjusted	0.20	0.21	0.22	0.18	0.21
Sample size	821,631	821,631	821,631	555,168	542,697

Note: Robust standard errors in (.). We exclude families having twins at the first birth. 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: 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				
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)
Family size measures				
More than two children	0.707	0.484	0.490	0.425
	(0.45)	(0.50)	(0.50)	(0.49)
Complete family size	3.046	2.606	2.612	2.509
	(0.91)	(0.73)	(0.73)	(0.68)
Instrument for fertility				
Twins at 2nd birth	0.0071	0.0069	0.0067	0.0061
	(0.08)	(0.08)	(0.08)	(0.08)
Covariates				
Urban (place of birth)	0.340	0.342	0.338	0.338
5-year average taxable income per capita in district (thousands)	730.0	730.8	729.6	729.2
Mother's age at 2nd 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+	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+	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
Gestation duration of 2nd birth (weeks)	39.66	39.61	39.63	39.59
Mean birth weight of 2nd birth (kg)	3.231	3.339	3.219	3.320

Table 4: Regressions of Birth Spacing (Measured in Months) between the First Two Births

	First Child Born in 1978-1984	First Child Born in 1978-1979	
		All	Next Sibling Born by 1985
	(1)	(2)	(3)
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	6.524	5.251	2.868
capita in village (thousands))	(0.145)	(0.261)	(0.157)
Sample size	820,162	238,554	228,753
R-squared adjusted	0.049	0.044	0.063

Note: Robust standard errors in (.). Regressions in this table include the same set of covariates as Table A2.

Table 5: Regressions of Twining at the 2nd Birth on Family Backgrounds

	(1)	(2)	(3)	(4)
Birthyear of 2nd birth	Yes	Yes	Yes	Yes
	[0.0000]	[0.0000]	[0.0000]	[0.0002]
Urban		0.00058	0.00059	0.00048
		(0.00024)	(0.00024)	(0.00024)
Ln(taxable income per		-0.00010	-0.00056	-0.00094
capita in district of birth)		(0.00051)	(0.00052)	(0.00053)
Parental Education			Yes	Yes
			[0.059]	[0.771]
Mother's age at 2nd birth				Yes
				[0.0000]
R-squred adjusted	0.00015	0.00016	0.00017	0.00024

Note: Robust standard errors in (.) and p-values for joint hypothesis tests in [.]. N=821,631.

Table 6: OLS and 2SLS Estimates for Regressions of High School Completion

	First-Born Girls				First-Born Boys			
Dependent Variable	OLS	2SLS	2SLS	2SLS	OLS	2SLS	2SLS	2SLS
= HS Completion	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	Mediator = $Morethan2$							
$Boy2nd$	-0.0014	-0.0057	-0.0602	-0.0628	-0.0050	-0.0046	-0.0061	-0.0076
	(0.0013)	(0.0039)	(0.0227)	(0.0227)	(0.0012)	(0.0015)	(0.0093)	(0.0094)
$Morethan2$	-0.0116	-0.0310	-0.1042	-0.1082	-0.0276	-0.0209	-0.0337	-0.0372
	(0.0015)	(0.0165)	(0.0341)	(0.0341)	(0.0013)	(0.0133)	(0.0202)	(0.0204)
$Morethan2 \times Boy2nd$			0.0790	0.0829			0.0015	0.0047
			(0.0313)	(0.0313)			(0.0187)	(0.0188)
Gestation period at				-0.0004				-0.0007
2nd birth (weeks)				(0.0005)				(0.0004)
	Mediator = $Sibsize$							
$Boy2nd$	-0.0044	-0.0116	-0.1031	-0.1098	-0.0057	-0.0050	0.0009	-0.0060
	(0.0013)	(0.0053)	(0.0469)	(0.0473)	(0.0012)	(0.0016)	(0.0368)	(0.0374)
$Sibsize$	-0.0127	-0.0293	-0.0527	-0.0550	-0.0242	-0.0169	-0.0252	-0.0282
	(0.0009)	(0.0118)	(0.0171)	(0.0173)	(0.0009)	(0.0105)	(0.0153)	(0.0155)
$Sibsize \times Boy2nd$			0.0312	0.0334			-0.0027	0.0000
			(0.0151)	(0.0153)			(0.0141)	(0.0143)
Gestation period at				-0.0005				-0.0008
2nd birth (weeks)				(0.0005)				(0.0004)

Note: We include 401,900 first-born girls and 419,731 first-born boys. Additional covariates include parental age, mother's age at second birth, subject's age, birthplace, urban dummy, parental education dummies, and logarithm of taxable income per capita in district of birth. We include interactions between the twins instrument and sibling sex as an additional instrument for family size also in the model without interactions between family size and sibling sex.

Table 7: Decomposing the Average Total Effect of *Boy2nd* on the First Child's High School Completion

	First-Born Girls				First-Born Boys			
	OLS	2SLS	2SLS Interact	2SLS Add GA	OLS	2SLS	2SLS Interact	2SLS Add GA
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Average Total Effect	0.0012	0.0012	0.0013	0.0015	-0.0032	-0.0032	-0.0032	-0.0032
	(0.0013)	(0.0013)	(0.0013)	(0.0013)	(0.0012)	(0.0012)	(0.0012)	(0.0012)
	Mediator = $Morethan2$							
(1) Average Indirect Effect	0.0026	0.0069	0.0232	0.0241	0.0018	0.0014	0.0022	0.0024
	(0.0003)	(0.0037)	(0.0076)	(0.0076)	(0.0001)	(0.0009)	(0.0013)	(0.0013)
(2) Average Direct Effect	-0.0014	-0.0057	-0.0220	-0.0227	-0.0050	-0.0046	-0.0054	-0.0056
	(0.0013)	(0.0039)	(0.0077)	(0.0077)	(0.0012)	(0.0015)	(0.0018)	(0.0018)
Controlled Direct Effect	-0.0014	-0.0057	-0.0135	-0.0138	-0.0050	-0.0046	-0.0054	-0.0054
	(0.0013)	(0.0039)	(0.0044)	(0.0044)	(0.0012)	(0.0015)	(0.0014)	(0.0014)
	Mediator = $Sibsize$							
(1) Average Indirect Effect	0.0056	0.0129	0.0232	0.0242	0.0025	0.0017	0.0026	0.0029
	(0.0004)	(0.0052)	(0.0075)	(0.0076)	0.0001	(0.0011)	(0.0016)	(0.0016)
(2) Average Direct Effect	-0.0044	-0.0116	-0.0219	-0.0227	-0.0057	-0.0050	-0.0058	-0.0061
	(0.0013)	(0.0053)	(0.0076)	(0.0076)	0.0012	(0.0016)	(0.0020)	(0.0020)
Controlled Direct Effect	-0.0044	-0.0116	-0.0153	-0.0156	-0.0057	-0.0050	-0.0060	-0.0061
	(0.0013)	(0.0053)	(0.0045)	(0.0045)	0.0012	(0.0016)	(0.0015)	(0.0015)

Note: "GA" stands for the length of the gestation period in weeks for the 2nd birth. Robust standard errors in (.). The controlled direct effect is evaluated at the mean of the fertility variable. For the list of control and instrumental variables, see Table 6.

Table 8: Decomposing the Average Total Effect of *Boy2nd* on the First Child's University Admission

	First-Born Girls				First-Born Boys			
	OLS	2SLS	2SLS Interact	2SLS Add GA	OLS	2SLS	2SLS Interact	2SLS Add GA
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Average Total Effect	0.0019	0.0019	0.0019	0.0020	0.0005	0.0005	0.0005	0.0006
	(0.0011)	(0.0011)	(0.0011)	(0.0012)	(0.0011)	(0.0011)	(0.0011)	(0.0011)
	Mediator = $Morethan2$							
(1) Average Indirect Effect	0.0024	0.0053	0.0173	0.0174	0.0012	0.0008	0.0011	0.0018
	(0.0003)	(0.0033)	(0.0068)	(0.0068)	(0.0001)	(0.0008)	(0.0009)	(0.0011)
(2) Average Direct Effect	-0.0005	-0.0034	-0.0153	-0.0154	-0.0007	-0.0004	-0.0006	-0.0012
	(0.0012)	(0.0035)	(0.0069)	(0.0069)	(0.0011)	(0.0013)	(0.0014)	(0.0016)
Controlled Direct Effect	-0.0005	-0.0034	-0.0092	-0.0091	-0.0007	-0.0004	-0.0006	-0.0011
	(0.0012)	(0.0035)	(0.0040)	(0.0039)	(0.0011)	(0.0013)	(0.0014)	(0.0012)
-0.000628	Mediator = $Sibsize$							0.0014149
(1) Average Indirect Effect	0.0049	0.0049	0.0172	0.0175	0.0017	0.0011	0.0019	0.0021
	(0.0003)	(0.0003)	(0.0068)	(0.0068)	(0.0001)	(0.0009)	(0.0014)	(0.0014)
(2) Average Direct Effect	-0.0030	-0.0030	-0.0153	-0.0154	-0.0012	-0.0006	-0.0014	-0.0015
	(0.0012)	(0.0012)	(0.0069)	(0.0069)	(0.0011)	(0.0014)	(0.0017)	(0.0017)
Controlled Direct Effect	-0.0030	-0.0030	-0.0106	-0.0106	-0.0012	-0.0006	-0.0015	-0.0015
	(0.0012)	(0.0012)	(0.0040)	(0.0040)	(0.0011)	(0.0014)	(0.0013)	(0.0013)

Note: Robust standard errors in (.). Same as Table 7.

Table 9: Heterogeneous Effects of *Boy2nd* on First-Born's Education (Divided by Sample Mean) by Maternal Education and Place of Birth

	First-Born Girls				First-Born Boys			
Effect Divided by Sample Mean	Univ (1)	HS (2)	Univ (3)	HS (4)	Univ (5)	HS (6)	Univ (7)	HS (8)
	Urban		Rural		Urban		Rural	
Average Total Effect	0.008	0.003	0.014	0.008	-0.008	-0.015	0.011	-0.012
	(0.010)	(0.008)	(0.009)	(0.007)	(0.011)	(0.008)	(0.009)	(0.007)
(1) Average Indirect Effect	0.147	0.097	0.040	0.100	0.001	0.002	0.024	0.020
	(0.045)	(0.035)	(0.063)	(0.050)	(0.007)	(0.005)	(0.013)	(0.009)
(2) Average Direct Effect	-0.139	-0.094	-0.026	-0.092	-0.008	-0.017	-0.014	-0.033
	(0.046)	(0.036)	(0.063)	(0.051)	(0.013)	(0.010)	(0.015)	(0.011)
Controlled Direct Effect	-0.082	-0.058	-0.017	-0.055	-0.012	-0.019	-0.007	-0.028
	(0.027)	(0.021)	(0.035)	(0.028)	(0.012)	(0.009)	(0.011)	(0.008)
	Mother JHS+		Less than JHS		Mother JHS+		Less than JHS	
Average Total Effect	-0.001	0.000	0.027	0.012	0.002	-0.009	0.006	-0.019
	(0.008)	(0.006)	(0.011)	(0.009)	(0.009)	(0.006)	(0.011)	(0.008)
(1) Average Indirect Effect	0.083	0.089	0.096	0.083	0.002	0.001	0.029	0.028
	(0.034)	(0.027)	(0.082)	(0.066)	(0.005)	(0.004)	(0.016)	(0.012)
(2) Average Direct Effect	-0.083	-0.088	-0.070	-0.070	-0.001	-0.010	-0.023	-0.046
	(0.035)	(0.028)	(0.083)	(0.066)	(0.010)	(0.007)	(0.020)	(0.015)
Controlled Direct Effect	-0.054	-0.055	-0.036	-0.042	-0.002	-0.011	-0.016	-0.040
	(0.021)	(0.016)	(0.046)	(0.037)	(0.009)	(0.007)	(0.014)	(0.011)

Note: Estimation is based on mediating variable $Morethan2$. Robust standard errors in fraction of sample mean are in (.). Covariates are the same as columns (4)(8) of Table 6. "HS" indicates high school completion, "Univ" university attainment, and "JHS" junior high school.

Table A1: Variable Mean for Families with One or More Children

	First Child Born in 1978-1984	First Child 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
Covariates:			
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
Mothers' highest degree			
College/professional degree+	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+	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: Testing Randomness of *Boy2nd*,
Regressions of *Boy2nd* on Family Backgrounds

	First Child Born in 1978-1984	First Child Born in 1978-1979	
		All	Next Sibling Born by 1985
	(1)	(2)	(3)
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+	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+	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

Note: Robust standard errors in (.).

Table A3: 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)
	Dependent Variable = $Morethan2$					
$Boy2nd$	-0.221	-0.223	-0.223	-0.064	-0.065	-0.064
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
$Twin2nd$	0.446	0.331	0.335	0.578	0.548	0.548
	(0.005)	(0.006)	(0.006)	(0.004)	(0.006)	(0.006)
$Twin2nd \times Boy2nd$		0.225	0.224		0.062	0.062
		(0.008)	(0.008)		(0.008)	(0.008)
Gestation period at 2nd birth (weeks)			0.0020			-0.0007
			(0.0005)			(0.0005)
	Dependent Variable = $Sibsize$					
$Boy2nd$	-0.437	-0.437	-0.436	-0.102	-0.102	-0.101
	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)
$Twin2nd$	0.645	0.650	0.651	0.733	0.724	0.717
	(0.012)	(0.020)	(0.020)	(0.010)	(0.015)	(0.015)
$Twin2nd \times Boy2nd$		-0.009	-0.007		0.018	0.020
		(0.024)	(0.024)		(0.021)	(0.021)
Gestation period at 2nd birth (weeks)			0.0004			-0.0033
			(0.0009)			(0.0008)

Note: Robust standard errors in (.). Additional covariates include parental age, mother's age at second birth, subject's age, birthplace, urban dummy, parental education dummies, and logarithm of taxable income per capita in district of birth.

Table A4: First-Stage Estimates for Interaction between Family Size and Sibling Sex, Instrumented by Interaction between Twinning at Second Birth and Silbing Gender

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

Note: Same as Table A3.

Table A5: 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
Covariates				
Urban (place of birth)	0.341	0.358	0.338	0.366
5-year average taxable income per capita in district (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+	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+	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
Gestation duration of 2nd birth (weeks)	39.64	38.12	39.62	38.10
Mean birth weight of 2nd birth (kg)	3.292	2.541	3.276	2.527