

Sibling size and investment in children's education: an asian instrument

Jungmin Lee

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Abstract This study estimates the trade-off between child quantity and quality by exploiting exogenous variation in fertility under son preferences. Under son preferences, both sibling size and fertility timing are determined depending on the first child's gender, which is random as long as parents do not abort girls at their first childbearing. For the sample South Korean households, I find strong evidence of unobserved heterogeneity in preferences for child quantity and quality across households. The trade-off is not as strong as observed cross-sectional relationships would suggest. However, even after controlling for unobserved heterogeneity, a greater number of siblings have adverse effects on per-child investment in education, in particular, when fertility is high.

Keywords Fertility · Education · Child quality

JEL Classification D13 · J13 · O12

1 Introduction

High fertility has been pointed out to be a major obstacle to human capital accumulation in developing countries, such as those in Africa and South Asia. The World Bank (1994) emphasizes the fundamental role of active population policy for poverty reduction and economic growth. From the perspective of endogenous economic growth models, the underlying rationale is straightforward: The quality of

Responsible editor: Junsen Zhang

J. Lee

Department of Economics, Sam M. Walton College of Business, University of Arkansas,
Fayetteville, AR 72701, USA

J. Lee (✉)

Institute for the Study of Labor (IZA), Bonn, Germany
e-mail: jlee@walton.uark.edu

the labor force, rather than the quantity, is essential for economic development, and given the scarcity of resources, a country can educate each worker better when there are fewer workers. Indeed, we find that fertility and per-child investment in education are significantly and negatively correlated across countries. This suggests that slowing down population growth should be a high priority in poor countries. Successful examples are East Asian countries, such as South Korea, Hong Kong, and China; these countries have dramatically restrained population growth and simultaneously increased educational investment over the past decades. At the household level, many studies, mostly in sociology, have also found that the number of siblings exerts a negative effect on each child's educational attainments such as grade completion and test scores. In one extensive study across various samples, Blake (1989) finds the "dilution effect": More siblings dilute a child's allocation of parental resources.¹

Economists have cast doubt on these empirical findings. The main reason is that both child quantity and quality should be simultaneously determined by parents.² Therefore, an observed inverse relationship may be spurious due to unobserved heterogeneity across households. Suppose that parents differ in preferences for child quality, then those who are more concerned about child quality would have fewer children to educate each better rather than have many mediocre ones. To avoid the simultaneity bias, Rosenzweig and Wolpin (1980) and Rosenzweig and Schultz (1987) exploited exogenous variations in fertility due to twinning.

Along with economics literature, this study attempts to address the problem of unobserved heterogeneity and consistently estimate the trade-off between child quantity and quality for sample South Korean households. This country provides an interesting context in which to examine the current issue. As one of the "Asian Tigers," it has been a showcase for developing countries on a fast track to economic development. Moreover, it has been often argued that a major factor behind its economic success was its population policy. The government established the Planned Parenthood Federation of Korea in 1961, which enforced four 5-year national plans from 1962 to 1981, implementing various policies to restrain population growth, such as distributing contraceptive measures, deregulating abortions, and campaigning for small-size family norms. Because many countries are trying to replicate this success story, it is essential to know, rather than recognize a simple correlation, the extent to which lowered fertility causally promoted investment in children's education.

There are two methodological contributions of this study to the literature. First, I use parents' monetary investment in children's education as a measure of child

¹ Interestingly, Gomes (1984) found for a household sample from Kenya that children from a larger family are more likely to complete grades. The reason is that parents in Kenya control their eldest child's earnings and younger children benefit from this extra source of family resources. This suggests that the relationship between child quantity and quality take different forms across different cultures. Cross-country comparative studies are demanded.

² On the other hand, economists have also tried to develop theoretical explanations for the observed trade-off. The major novelty here is that, even without assuming unusual substitutability between child quantity and quality in preferences, the trade-off may exist due to their interaction due to budget constraints (Becker and Lewis 1973).

quality, instead of the traditional measures such as children's schooling, test scores, or earnings. The reason is twofold; first, monetary investment should be as legitimate for a measure of child quality as the traditional variables because it is an important input for educational production. Second, because the economic theory of the trade-off between child quantity and quality is about parents' choices, it should be more appropriate to directly look at parental decisions to test this theory. The traditional measures are the outcomes of such parental decisions made when children are small and adolescent, and are realized far after parents decide to bear children. As such, they are likely to be tainted with various factors that might be independent of the parents' intention.

The second contribution is the use of a new instrumental variable (IV) for fertility—the first child's sex. It is well known that people in several Asian countries prefer sons to daughters. Son preferences lead to some particular patterns in fertility, which are useful for the current study. For example, consider that a couple with one child is thinking about whether to have another child or not. If parents prefer to have at least one son, and if the first child is not a son, they are more likely to try to have another child. Therefore, in a society with son preferences, the first child's sex should be a good predictor for the probability of having a second child or the total number of children. This simple example shows that the first child's sex is a good candidate for the IV for the actual number of children, as long as it is uncorrelated with the dependent variable, that is, parental investment in children's education in this study.

The empirical strategy of this study is motivated by Iacovou (2001) and Angrist and Evans (1998), but innovates upon their approach.³ They exploited preferences for a balanced sex composition of children in Western countries (the United States and the United Kingdom). Both studies found that parents with same-sex children are more likely to have additional children. A technical limitation of the studies is that they had to focus on the marginal effect of a third child. On the other hand, I can estimate the marginal effect of a second child, which is more important in low-fertility countries like South Korea and most developed countries. I also extend the previous studies by fully exploiting differences in the mothers' age profile of fertility (fertility timing as well as number of children over age from 21 to 45 by the first child's sex. The resulting multiple IVs allow us to estimate the nonlinear effects of child quantity on quality.

2 Data and descriptive statistics

The data used in this study are from the Korean Household Panel Study (KHPS). The KHPS is the first panel study on South Korean households conducted by a private institution, Daewoo Economic Research Institute. Our sample spans the entire period over which the data are available between 1993 and 1998.

The sample is restricted by the following steps; first, the sample should be limited to those households with at least one child to use the first child's sex as an IV. This

³ There are a few recent papers using a similar empirical strategy. For example, see Chun and Oh (2002); Black et al. (2005); Conley and Glauber (2005); and Angrist et al. (2005).

does not cause any serious sample selection problem, as almost every household has at least one child at some point in the course of life. Next, I focus on those households where all children are in schooling, that is, before college entrance and those with the first child under age 19. It is well known that college entrance is highly competitive in South Korea and drives up households' educational expenditure. Given that college quality is a major determinant for future earnings and social status, it is reasonable to assume that the educational expenditure for these young children represents parents' concerns about child quality. On the other hand, the expenditure for college students is determined by various factors other than parents' willingness to invest, such as their academic ability. I also exclude those with more than three children for homogeneity of the sample. This is because a major objective of this paper is to estimate the nonlinear effects of sibling size on educational investment. Having few observations with four children prevents us from estimating the marginal effect of a fourth child. The sample selection problem from this restriction should be ignorable, as only 38 observations are excluded. Finally, those households with missing or unreliable information about crucial variables are deleted. The final sample is an unbalanced panel data set consisting of 5,180 observations on 1,663 households over 6 years.

Table 1 shows descriptive statistics. The first notable thing is that, although the country is a rapidly developing country, its fertility is quite low and is comparable to that of the United States. The low fertility rate is not due to the sample selectivity. It is actually consistent with the national average number of births per woman aged 15 to 49 that was 1.4 in 1998 according to national statistics based on birth registration. One thing to be noted in this study is that the low fertility rate is inconsistent with the conventional wisdom that son preferences tend to increase fertility because parents continue to have children until they achieve their desirable number of sons. The coexistence of son preferences and low fertility indicates widespread sex-selective abortions (Park and Cho 1995). I will discuss this later in detail.

The most crucial variable is parental investment in children's education, measured by total monthly household expenditure on such items as tuition, schoolbooks and supplies, reference books, and private tutoring. Among others, it is notable that private tutoring accounts for 28% of the total education expenditure and 7.3% of the household income.⁴ A large amount of spending on private tutoring might be surprising to readers who are not familiar with the Korean education system. First of all, it should be noted that this is not specific to my sample. In fact, the sample average is quite consistent with national statistics showing that private tutoring expenses are 7.4 to 9% of household income. Kim and Lee (2002) also reported that private tutoring is widespread across different income groups. For example, they found that more than 70% of elementary school students and about half of middle and high school students took some private tutoring in 1997. Besides the fact that private tutoring accounts for a substantial portion of total education expenditures, it

⁴ School supplies and reference books are likely to account for most of the rest of total expenses. In particular, most students study with some "unrequired" reference books that are quite expensive. Tuition and other school fees are relatively small. For example, in 2005, annual high-school tuition and fees were around US \$1,500.

Table 1 Descriptive statistics

Year	1993	1994	1995	1996	1997	1998
Fathers						
Age	36.42 (4.26)	36.87 (4.41)	37.11 (4.40)	37.73 (4.59)	38.22 (4.63)	38.82 (4.71)
Schooling	12.46 (2.78)	12.44 (2.64)	12.64 (2.60)	12.66 (2.62)	12.73 (2.63)	12.78 (2.58)
Mothers						
Age	33.12 (3.65)	33.53 (3.81)	33.79 (3.86)	34.38 (4.03)	34.91 (4.13)	35.46 (4.25)
Schooling	11.46 (2.43)	11.54 (2.32)	11.76 (2.33)	11.83 (2.40)	11.93 (2.35)	12.02 (2.26)
Households						
Seoul	0.31 (0.46)	0.33 (0.47)	0.32 (0.47)	0.31 (0.46)	0.32 (0.47)	0.28 (0.45)
Large Cities	0.28 (0.45)	0.27 (0.44)	0.25 (0.43)	0.26 (0.44)	0.25 (0.44)	0.28 (0.45)
Grandparents	0.07 (0.25)	0.08 (0.28)	0.07 (0.26)	0.07 (0.25)	0.08 (0.27)	0.09 (0.29)
Mean Child Age	7.09 (3.63)	7.37 (3.80)	7.45 (3.78)	7.97 (3.89)	8.37 (3.98)	8.81 (4.04)
Middle School	0.08 (0.21)	0.11 (0.25)	0.11 (0.25)	0.13 (0.26)	0.13 (0.27)	0.14 (0.27)
High School	0.02 (0.12)	0.03 (0.13)	0.03 (0.15)	0.05 (0.16)	0.07 (0.20)	0.09 (0.24)
Number of Children	1.91 (0.52)	1.93 (0.51)	1.92 (0.50)	1.93 (0.48)	1.92 (0.50)	1.95 (0.50)
Monthly Investment in Education	0.18 (0.12)	0.20 (0.14)	0.24 (0.16)	0.27 (0.18)	0.29 (0.18)	0.24 (0.16)
Monthly Household Income	2.00 (1.16)	2.13 (1.15)	2.35 (1.19)	2.59 (1.32)	2.60 (1.15)	2.12 (1.14)
N=	1,136	967	874	849	800	554

Monthly investment in education and monthly household income are denoted in 100,000 Korean won and real value

is important to note that it also explains about 50% of the total variation in educational expenditure across households. Tuition and fees do not differ between public and private schools due to government policies such as universal elementary education and the so-called equalization policy for middle and high schools. Except for colleges and universities, private schools are virtually equivalent to public schools in terms of curricula as well as costs. Kim and Lee (2002) claimed that the prevalence of private tutoring is a market response to the equity-centered education policies. Some parents are willing to invest more in their children, but there is no opportunity within the formal educational system. Thus, the market for private tutoring has evolved to address parents' excessive demand for children's education.

It is also notable that real educational expenses consistently increased over time, except in 1998 when the economy severely suffered in the midst of the Asian financial crisis. And the proportion of education investment in total household income steadily increases from 7.7% in 1993 to 12.7% in 1997 and, despite the crisis, to 13.1% in 1998. This suggests that children's education not only accounts

for a large portion of household expenditure but also takes a high priority in household resource allocation.⁵

3 First child's sex as an instrument for fertility

3.1 Sibling number and fertility timing

Under son preferences, parity progression and fertility timing should depend not only on the gap between the desired and current number of total children but also on the gap between parents' preferred and present number of sons. For example, if a couple with one child desired to have one son but failed to have one until now, they should not only try to have another child but also should do so more quickly. Although the first child's sex can be used to predict the likelihood of having a second child given the mother's age, it is also a good predictor for the existence of a third child because those who had a daughter as their first child could have two daughters consecutively and, as a result, are more likely to have a third child. In general, those who had a girl at first are more likely to end up with more children at any given age.

The predictive power of the first child's sex, which is a critical qualification for the IV, depends not only upon the strength of son preferences but on the social norms about family size. Assuming that parents prefer a smaller number of children if other things are equal, we can derive the following results⁶; first, having a girl at first (hereinafter, *first girl*) leads to mothers at a given age having a larger number of children later. Second, the marginal effect of *first girl* on child quantity would increase as more parents in the population prefer sons or as parents have a larger desired number of sons. Lastly, the marginal effect is larger as parents minimize the number of children. These results suggest that *first girl* is likely to qualify for the IV for child quantity in the context of South Korea where son preferences are strong and, on the other hand, a small family is a social norm.

The above results are testable. Two tests are feasible with my sample. First, I tested for the existence of son preferences by estimating a parity progression model, which is widely used in the demography literature. The hypothesis is that, if there were son preferences, those who had more daughters in earlier phases of fertility should end up with a larger number of children. It is important to note that the parity progression model must be tested for those who have completed fertility because households' decision about further childbearing should be fully observed (not right-censored). Following the convention in demography, I assume that mothers aged 40 and older are no longer fertile. I estimate a Logit model where the dependent variable is an indicator of the parity progression (e.g., whether parents have more than one child or not). The key explanatory variable is the sex composition of children under a certain birth order.

⁵ Why South Korean parents are strongly concerned about children's education is an important question but beyond the scope of this paper. It might be because of strong intergenerational ties or increasing demand for skilled workers. Refer to Seth (2002) for a historical perspective on the origin of "education fever" in South Korea.

⁶ An earlier version of this paper proved these results and is available from the author upon request.

Table 2 shows results. The first column shows that *first girl* increases the probability of having a second child (9.8%). In the second column, having two same-sex children (*same sex*) increases the probability of having three or more children (9.9%). This positive effect turns out to be entirely due to a large positive effect of having two girls (*GG*), about 34.8%. In fact, having two boys (*BB*) has a negative effect on fertility, implying that parents with two sons are more likely to stop childbearing as compared to the reference group of parents of one boy and one girl. South Korean parents do not have preferences for a balanced sex mix of children.

As a second test for son preferences, I estimate a hazard model of fertility timing. The hypothesis in this paper is that, given a time (biological) constraint for childbearing, parents should advance the timing of the second birth after the first daughter more than they do after the first son. By the same logic, they should try to have a third child sooner if they had two daughters. The dependent variable is the time interval between two consecutive births. An advantage of the hazard model over the parity progression model is that we do not have to restrict the sample. Specifically, I estimate a Weibull hazard model, taking into account right censoring of the birth interval. Only time-invariant variables, such as parents’ education levels and the mother’s age at the first birth, are included to avoid any bias that arises because explanatory variables change during the interval.

Relative hazard ratios are calculated in Table 3. As expected, *first girl* significantly increases the hazard rate of having a second child by at least 48%. As before, any positive effect of *same sex* disappears once we include *BB* and *GG* separately, as with the results of parity progression model. The coefficient of *BB* is estimated to be less than one, which means that having two boys decreases the hazard rate of having a third child. This implies that parents are likely to delay additional childbearing—probably permanently—if they have two boys. On the other hand, *GG* significantly advances the third childbirth.

In summation, the instrument variable, *first girl*, is strongly correlated with child quantity dependent on the mother’s age. In other words, married women have

Table 2 Logit models of parity progression

Dependent variable	1[n>1]	1[n>2]	1[n>2]
First girl	0.098 (0.019)	0.286 (0.035)	0.036 (0.046)
Second girl		0.251 (0.033)	
Same sex (BB or GG)		0.099 (0.035)	
BB			-0.151 (0.044)
GG			0.349 (0.045)
Pseudo R-squared	0.158	0.151	0.151
N	4,259	3,373	3,373

Robust standard errors are displayed in parentheses. Equations include mother age, mother age squared, mother age at first birth, grandparents, Seoul, and large cities.

Table 3 Weibull hazard models of childbirth spacing

Dependent variable	AGE1–AGE2	AGE2–AGE3	AGE2–AGE3
First girl	1.476 (0.111)	2.559 (0.504)	0.704 (0.211)
Second girl		3.452 (0.693)	
Same sex (BB or GG)		2.450 (0.479)	
BB			0.710 (0.227)
GG			8.457 (1.972)
Pseudo <i>R</i> -squared			
<i>N</i>	7,149	5,363	5,363

Hazard ratios are presented. Robust standard errors are displayed in parentheses. Equations include mother age, mother age at first birth, and mother education. AGE_{*i*} represents the age of child of birth order *i*.

different age profiles of fertility depending on the first child's sex. It should be able to serve as an instrument for child quantity and, given the significance of the instrument in determining child quantity, most likely pass the chronic problem of "weak instrument" (Staiger and Stock 1997; Bound et al. 1995).

3.2 Sex-selective abortion

One might ask why I use only the first child's sex as an IV. The question seems natural because, for example, the sex composition of the first two children should predict child quantity better than the first child's sex alone for those who have more than one child. If parents had two daughters consecutively, they are more likely to have another child than those who had two sons or those who had a son and a daughter.

However, the prevalence of sex-selective abortions mentioned before prevents us from utilizing the second child's sex as an additional IV. It is well known that sex-selective abortions were very common in South Korea, in particular, during the 1990s including the sample period between 1993 and 1998 (Park and Cho 1995). To illustrate this, Table 4 shows the sex ratio at birth (the number of boys divided by the number of girls times 100) by birth order between 1981 and 2000. If children's sex was randomly assigned by nature and the infant mortality rate was constant over birth order and between sexes, the sex ratio should have stayed around the normal level (about 106). However, we find that the sex ratio is rapidly imbalanced, as the parity is higher. The increasing sex ratio implies that parents not only prefer to have sons, but also prefer to have a small number of children. Thus, female births after a certain number of children are suppressed.

It is surprising that the imbalanced sex ratios were found even at a very low birth order, two, since the mid-1980s. This suggests that even the second child's sex is not randomly assigned by nature but chosen by the parents to some extent. As a result, the second child's sex might also be correlated with investment in education, depending on the unobservable (parents' preferences). Those who strongly prefer a son and, therefore, are more willing to abort girls might invest more in children's

Table 4 Sex ratio at birth by birth order

Birth order	One	Two	Three	Four and higher	Total
1981	106.3	106.7	107.1	112.9	107.2
1982	105.4	106.0	109.2	113.6	106.8
1983	105.8	106.2	111.8	120.0	107.4
1984	106.1	107.2	116.9	128.1	108.3
1985	106.0	107.8	129.2	146.8	109.4
1986	107.3	111.2	138.6	149.9	111.7
1987	104.7	109.1	134.9	148.8	108.8
1988	107.2	113.2	164.4	181.7	113.3
1989	104.1	112.4	181.4	198.7	111.7
1990	108.5	117.0	188.8	209.2	116.5
1991	105.7	112.4	179.5	193.5	112.4
1992	106.2	112.4	191.1	212.1	113.6
1993	106.4	114.7	202.1	234.8	115.3
1994	106.0	114.1	202.2	223.5	115.2
1995	105.8	111.7	177.2	203.7	113.2
1996	105.3	109.8	164.0	183.4	111.6
1997	105.1	106.3	133.5	153.5	108.2
1998	105.9	108.0	144.7	153.2	110.2
1999	105.6	107.6	141.8	153.8	109.6
2000	106.2	107.4	141.7	166.9	110.2

Korea National Statistical Office, vital statistics. The world average is about 106.

education because they have more “favored” children. On the other hand, the sex ratio for the birth order one remained around the normal level during the 1990s. Parents seem to leave at least the sex of the first child to natural selection. The first child's sex is likely to be orthogonal to parents' preferences.

3.3 Distinguishing prenatal vs postnatal son preferences

Now, I examine another requirement for consistent instrumental-variable estimation—the “validity” condition. In the context of this paper, it requires that the first child's sex should not directly affect parental investment in education, but only does so indirectly through the number of children. Readers might suspect that educational investment would depend upon the first child's sex, as the latter is correlated with the overall sex composition of children. That is, if the first child is a son, the household is likely to have more sons. It might be true under son preferences that those whose first child is a son might invest more in children's education just because they have more favored children.

It is, however, important to distinguish prenatal son preference from postnatal son preference. The first means that parents prefer to have sons rather than daughters, whereas the latter implies that they treat sons in more favorable ways after they are born. The particular patterns in childbearing and fertility timing that we previously observed are related to prenatal son preferences. What matters now for the validity condition is the existence of postnatal son preferences.

Although these two types of son preferences are likely to be closely related, in theory, they do not necessarily coexist. In fact, as Goodkind (1996) mentioned,

Table 5 Educational expenses by sibling size and sex composition

Children	Sons	N	Total education		Private tutoring		Others	
			Mean	Standard deviation	Mean	Standard deviation	Mean	Standard deviation
1	1	554	0.158	0.122	0.105	0.083	0.052	0.101
1	0	313	0.158	0.125	0.116	0.114	0.042	0.069
2	2	1,058	0.230	0.153	0.147	0.111	0.084	0.122
2	1	2,133	0.248	0.162	0.154	0.114	0.094	0.130
2	0	652	0.226	0.163	0.154	0.125	0.072	0.106
3	3	6	0.253	0.144	0.146	0.115	0.107	0.171
3	2	103	0.303	0.193	0.218	0.153	0.084	0.116
3	1	266	0.303	0.180	0.191	0.128	0.112	0.142
3	0	95	0.279	0.180	0.188	0.129	0.091	0.127

prenatal son preferences are likely to substitute for postnatal son preferences when sex-selective technologies are available and small-size families are preferred. The reason is that daughters who are not aborted even though abortion is available are more likely to be “wanted” children. This argument is likely to be true in South Korea where sex-selective abortions were widespread for the study period. Indeed, Goodkind (1996) pointed out that the substitution hypothesis holds true particularly in South Korea.

There seems to be no strong evidence for the existence of postnatal son preferences in the recent years after prenatal sex-selective technologies became available. The gender gap in the secondary school enrollment rate has decreased from 15% in 1970 to none in 1995. The probability of advancing from high school to college or university becomes even higher for females (75%) than for males (70%) in 1995. If there were substantial discrimination against girls in terms of educational expenditure, we should have observed significant gender differences in these educational outcomes.

One might think that even if there is no postnatal gender discrimination by parents, there might be pure educational cost differences between sexes. For example, if boys had more developmental problems, they would incur more costs. In this case, child gender would directly affect educational expenditure.⁷ To check this, in Table 5, I compare total educational expenditure with sibling size and sex composition. I find no significant gender difference in both private tutoring and other expenses, whereas total expenditure generally increases with the number of children. For example, there is no difference among households with two children between those with only sons and those with only daughters.⁸ It should be noted though that this simple comparison is intuitive; it does not account for unobserved heterogeneity. And some comparisons are not feasible due to a small number of observations.

⁷ Investment in education might depend on the sex composition of children due to pure cost differences in education across girls and boys. After controlling for sibling size, Lundberg and Rose (2004) find no significant difference in educational costs by gender in the United States.

⁸ The number of observations for each case is interesting. For example, there are only six observations with three sons and no daughters, which implies that most households stop further childbearing when they have two sons. Also the number of households with two daughters only is disproportionately small, which suggests that households are likely to have another child when they have two daughters.

The absence of postnatal son preferences might be puzzling because men face better earning opportunities than women do. An economic theory suggests that, in this case, sons should get more human capital investment than girls should. However, parents should decide investment based on their expectation about children's potential adult earnings (Davies and Zhang 1995). It should be noted that the gender inequality in the South Korean labor market has been decreasing in recent years. The unconditional gender earnings gap (the ratio of female to male earnings) was 47% in 1985, 53% in 1990, and 61% in 1997.⁹ Given this trend, the pure economic incentive to discriminate against daughters will not be as strong as it looks at present.

Even if there were strong postnatal preferences for sons and, as a result, if the IV estimates were biased, fortunately we can figure out the direction of bias. Postnatal preferences for sons suggest that the dependent variable (total investment in education) and the instrument (*first girl*) should be negatively correlated. Then the estimates will be underestimated. The true trade-off between child quantity and quality should be weaker than what the estimates imply.¹⁰

In addition, there is another case where the IV estimation could underestimate the true effect. As Hanushek (1992) pointed out, decreased child spacing acts like an increase in sibling number because smaller spacing means (1) lower probability of being in a small family and (2) less attention and resources from parents. As seen from the hazard model, under son preferences, child spacing between the first and second child tends to decrease if the first child is a daughter. However, it is tricky to control for birth spacing in my regression analysis because it is also endogenous. As a result, the marginal effects in absolute terms are likely to be underestimated. This omitted variable bias will again strengthen my conclusion that the trade-off between child quantity and quality is not as strong as what observed cross-sectional relationships suggest.

4 Regression analysis and results

I start by estimating a simple linear equation:

$$\ln I_{it} = \alpha n_{it} + X_{it}\beta + u_{1it}, \tag{1}$$

$$\text{where } n_{it} = X_{it}\gamma + Z_{it}\delta + u_{2it}. \tag{2}$$

The error term u_{1it} is assumed to be independent of all the control variables in X , such as mother's age, parents' years of schooling, the natural logarithm of monthly total household income, an indicator of living with grandparents, dummies for

⁹ The returns to college education are even higher for women; the college wage premium, measured by the average wage gap between high-school and college graduates as a proportion of high-school graduates' average wage, was 65% for women but only 44% for men in 1997. In addition, there is no gender gap in the employment rate for both high school and college graduates.

¹⁰ Kim and Lee (2002) found that parents invest slightly more in private tutoring for daughters. They interpret this finding as a result of the fact that girls are more likely to take private tutoring in music and arts, which tend to be more expensive. Their finding suggests that our estimates might be overestimated in absolute terms. To address this issue, we experimented with a different dependent variable, total investment in education except for private tutoring, which does not significantly change our results below.

residential location, and four yearly dummies. Because educational costs differ by levels of schooling, I include average age of children and its square term, as suggested by Browning (1992). The variable n represents the number of children in the household. Later, I will allow for the nonlinear effects of child quantity.

The variable I represents the monthly total investment in education for all children in the household. Formally, I equals $p_q(nq)$, where q is the average quality of children and p_q is the unit price of child quality (Becker and Lewis 1973). By the Slutsky theorem and product rule, the coefficient α represents:

$$\alpha = p \lim \frac{\partial \ln p_q n q}{\partial n} = p \lim \left(\frac{1}{n} + \frac{\partial q}{\partial n} \frac{1}{q} \right) = \frac{1}{E(n)} (1 + \varepsilon_{qn}), \quad (3)$$

where ε_{qn} represents the elasticity of child quality with respect to child quantity. Using the equation, I will recover the elasticity from the consistent estimate of α and the sample average of n .

Equation 2 specifies the first-stage regression in which fertility is endogenously determined by exogenous explanatory variables (X) and IVs (Z) excluded from the second-stage regression. Besides the first child's sex (*first girl*), there are four additional IVs, including its interactions with some exogenous variables in X .

First, I expect that the interaction of *first girl* and *grandparents* increases fertility, but does not directly affect total investment in education. When the household head is the eldest son, he is strongly asked to have at least one son to carry the responsibility for continuing family lineage. Thus, those households are more likely to have a second child once they had a girl at first. Unfortunately, there is no direct information about the birth order of household head. But, the indicator of living together with grandparents well represents whether the head is eldest or not. It requires a strong assumption that living arrangements should be exogenous to household decisions about investment in children. This assumption is valid in South Korea where living with old grandparents is socially enforced. But even if we drop this instrument, the results do not qualitatively change.

I also use the quadratic terms of the mother's age as it interacts with *first girl* as additional instruments. The rationale is that the age profile of fertility differs by the first child's sex, as the results from the hazard model imply that those who had a daughter at first are likely to have more children at certain age than those who had a son. Lastly, I include the interaction of the mother's education and *first girl* because lower-educated mothers put greater importance on traditional values, including son preferences. I assume that, after interaction with the first child's sex, the mother's education does not affect total investment in children's education.

Table 6 presents the estimation results for the linear specification. The ordinary least squares (OLS) estimates show that an additional child increases total investment by about 27%.¹¹ The OLS estimate implies the existence of the trade-

¹¹ Including the time-invariant random effect that is orthogonal to explanatory variables does not make any significant difference. Also, I ran a regression of educational investment directly on the first child's sex (or overall sex composition) with other controls. Even though the estimates are potentially biased, if the sex mattered in a significant way, I should have found some direct effects of the first child's sex (or sex composition). But I found no significant effect. Those variables are significant only if we lower the number of children.

off between child quantity and quality. The elasticity of the per-child investment with respect to fertility is negative (−0.49).

On the other hand, the IV estimates show that the OLS estimates are underestimated by 6 to 10 percentage points.¹² Unobserved heterogeneity accounts for some of the observed trade-off. The elasticity calculated based on the IV estimates ranges from −0.29 to −0.37, smaller in absolute terms than that based on the OLS estimates. But they are still negative.

The IVs in the first-stage regression have significant explanatory power. In column 2, *first girl* significantly increases the number of children. The partial *R*-squared amounts to 0.03, about 30% of the total *R*-squared. The *F*-statistic for the instrument is 174.8, much larger than the rule-of-thumb threshold of 10 (Bound et al. 1995). In column 4, the partial *R*-squared for multiple instruments is similar, and the *F*-statistic is 38.7. Having multiple instruments allows us to test for the validity of instruments by testing over-identifying restrictions. We find that the validity of instruments is not rejected (*p* value for Hansen’s *J*-statistic test is 0.72).

Table 7 shows some robustness checks. First, I use spending on private tutoring as a tightened dependent variable. Besides its quantitative importance, spending on private tutoring would be a part of educational expenses that parents can choose at their will. One may think that this narrow definition is more appropriate for the purpose of this paper. However, the results change little. An additional child increases total investment for children’s education only by 25%. After accounting for the endogeneity of child quantity, the marginal effect is about 48%. But the trade-off is still significant. Second, I check the robustness of our results to repeated observations. I restrict the sample to the first wave without attrition and run the same regressions. The results confirm the previous conclusion that, without accounting for the endogeneity, the trade-off is significantly exaggerated. In fact, the IV estimate suggests no trade-off; we cannot reject the hypothesis that the elasticity of child quality with respect to child quantity is zero (*p* value is 0.17). Lastly, I restrict the sample to those households in which mothers do not work in the labor market. One might think that, because high fertility encourages mothers to withdraw from the labor force and thereby decreases the household income, at least some part of the observed trade-off between fertility and per-child investment could be driven by the income effect (note, however, that the household income is already controlled). The results for nonworking mothers are very similar to my previous results.

As noted earlier, I relax the linearity of fertility in Eq. 1 by substituting *n* with two dummy variables; *D*₂=1 if there are exactly two children in the household and 0 otherwise; *D*₃=1 if there are exactly three children and 0 otherwise. Recall that the maximum number of children in the sample is three. Thus I estimate:

$$\begin{aligned} \ln I_{it} &= \alpha_2 D_{2,it} + \alpha_3 D_{3,it} + X_{it} \beta + u_{1it} \\ \text{where } D_{2,it} &= X_{it} \gamma_2 + Z_{it} \delta_2 + u_{2it} \\ D_{3,it} &= X_{it} \gamma_3 + Z_{it} \delta_3 + u_{3it}. \end{aligned} \tag{4}$$

¹² The IV estimates in this study are interpreted as local average treatment effects (Imbens and Angrist 1994). That is, my results show the marginal effects of the number of siblings for those who would change their childbearing decisions depending upon the first child’s sex. The monotonicity assumption is likely to hold for the current study because defiers (those with “daughter preferences”) are rare.

Table 6 Linear models of investment in children’s education

Regression stage Dependent variable	1	2	3	4	5
	OLS ln <i>I</i>	IV First <i>n</i>	Second ln <i>I</i>	IV First <i>n</i>	Second ln <i>I</i>
Number of children	0.267 (0.022)		0.369 (0.119)		0.328 (0.088)
Household income	0.422 (0.026)	0.061 (0.016)	0.416 (0.027)	0.060 (0.016)	0.419 (0.022)
Seoul	0.075 (0.025)	-0.118 (0.016)	0.088 (0.028)	-0.116 (0.016)	0.083 (0.023)
Large cities	0.044 (0.025)	-0.032 (0.016)	0.047 (0.025)	-0.031 (0.016)	0.046 (0.020)
Grandparents	0.001 (0.040)	0.080 (0.025)	-0.006 (0.040)	0.029 (0.033)	-0.003 (0.032)
Father education	0.013 (0.005)	0.002 (0.003)	0.013 (0.005)	0.002 (0.003)	0.013 (0.004)
Mother education	0.031 (0.006)	-0.027 (0.004)	0.034 (0.007)	-0.022 (0.005)	0.033 (0.006)
Mother age	0.051 (0.040)	0.141 (0.026)	0.038 (0.042)	0.160 (0.035)	0.044 (0.035)
Mother age squared	-0.001 (0.001)	-0.002 (0.0004)	0.000 (0.001)	-0.002 (0.001)	-0.001 (0.000)
Mean child age	0.278 (0.020)	0.099 (0.012)	0.268 (0.024)	0.098 (0.012)	0.272 (0.020)
Mean child age squared	-0.013 (0.001)	-0.008 (0.001)	-0.012 (0.002)	-0.008 (0.001)	-0.012 (0.001)
Middle school	0.262 (0.065)	0.461 (0.053)	0.214 (0.085)	0.460 (0.052)	0.233 (0.076)
High school	0.641 (0.127)	0.881 (0.093)	0.551 (0.167)	0.880 (0.093)	0.587 (0.142)
Year 1994	0.033 (0.023)	0.010 (0.021)	0.032 (0.023)	0.010 (0.021)	0.033 (0.026)
Year 1995	0.142 (0.026)	-0.013 (0.022)	0.144 (0.027)	-0.013 (0.022)	0.143 (0.027)
Year 1996	0.187 (0.026)	-0.012 (0.022)	0.188 (0.026)	-0.011 (0.022)	0.187 (0.027)
Year 1997	0.252 (0.028)	-0.032 (0.023)	0.255 (0.028)	-0.030 (0.023)	0.254 (0.028)
Year 1998	0.038 (0.033)	0.017 (0.025)	0.037 (0.033)	0.019 (0.025)	0.037 (0.033)
First girl		0.177 (0.013)		0.573 (0.789)	
× grandparents				0.125 (0.051)	
× mother education				-0.013 (0.006)	
× mother age				-0.022 (0.046)	
× mother age squared				0.000 (0.001)	
Constant	-5.483 (0.661)	-0.701 (0.437)	-5.439 (0.665)	-1.022 (0.578)	-5.456 (0.560)
<i>R</i> -squared	0.424	0.104	0.420	0.106	0.423
Partial <i>R</i> -squared		0.033		0.036	
<i>F</i> test		174.8		38.67	
(<i>p</i> value)		(<i>p</i> <0.01)		(<i>p</i> <0.01)	

Table 6 (continued)

	1	2	3	4	5
Regression stage	OLS	IV		IV	
Dependent variable	$\ln I$	First n	Second $\ln I$	First n	Second $\ln I$
Hansen J -statistic (p value)					2.09 ($p=0.72$)

$N=5,180$. Robust standard errors, calculated by clustering for households, are displayed in parentheses.

The coefficients, α_2 and α_3 , represent the marginal effects of the second and third child, respectively. Now that there are two endogenous explanatory variables, I include at least two IVs for identification. Table 8 presents results. Again, the OLS estimates imply the existence of the trade-off. I calculate the marginal effects on per-child investment from the coefficients by taking $(1 + \alpha_n)/n$ where $n=2$ or 3 . For the households with two children, per-child investment is only 67.7% of the amount of investment that a child could have received if there was only one child. For those with three children, it is much smaller, 49.4% of that of the reference group.

However, according to the IV estimates, per-child investment for the households with two children is 74.6%, and that for those with three children is 57.6% of what a child could have received if he or she were an only child. The marginal effects of child quantity on per-child investment are statistically different from zero at the 1% significance level. The results suggest that the trade-off exists and, as the number of children increases, the trade-off becomes stronger. However, the adverse effects of fertility on per-child investment are not as large as what the OLS estimates suggest. Remember that, if postnatal son preferences existed, even the IV estimates should exaggerate the trade-off. Thus, the true trade-off is likely to be much weaker than what the IV estimates suggest.

Some results in the first-stage regressions are worth noting. As expected, the interaction of the terms *first girl* and *grandparents* have significant positive coefficients. The results are consistent with the prior expectation that those households whose head is the first son have the familial responsibility of having at least one son. Unlike Table 6, the interaction terms with *mother age* and *mother age squared* are significant.

Table 7 Robustness checks

	OLS	IV
Private tutoring ($N=4,935$)	0.249 (0.026)	0.477 (0.141)
1993 sample ($N=1,136$)	0.196 (0.033)	0.704 (0.211)
Nonworking mothers ($N=3,487$)	0.266 (0.026)	0.351 (0.143)

The results for the coefficient on the number of children are only reported. The same set of control variables is included. Robust standard errors, calculated by clustering for households, are in parentheses. The sample size for the private tutoring estimation is smaller because there are some observations with zero spending on private tutoring.

Table 8 Nonlinear models of investment in children's education

	1	2	3	4
Regression stage	OLS	IV		
Dependent variable	$\ln I$	First 1 D2	First 2 D3	Second $\ln I$
Two children	0.354 (0.028)			0.492 (0.350)
Three children	0.481 (0.043)			0.727 (0.271)
Household income	0.423 (0.026)	0.003 (0.015)	0.029 (0.009)	0.416 (0.028)
Seoul	0.071 (0.025)	0.004 (0.014)	-0.060 (0.009)	0.086 (0.029)
Large cities	0.043 (0.025)	0.003 (0.015)	-0.017 (0.009)	0.046 (0.025)
Grandparents	0.001 (0.039)	-0.032 (0.030)	0.030 (0.019)	-0.008 (0.039)
Father education	0.013 (0.005)	-0.005 (0.003)	0.004 (0.002)	0.013 (0.005)
Mother education	0.030 (0.006)	-0.003 (0.004)	-0.009 (0.003)	0.033 (0.007)
Mother age	0.041 (0.040)	0.216 (0.032)	-0.028 (0.020)	0.024 (0.055)
Mother age squared	-0.0005 (0.001)	-0.003 (0.0005)	0.0005 (0.0003)	-0.0002 (0.001)
Mean Child Age	0.281 (0.020)	-0.001 (0.011)	0.049 (0.007)	0.269 (0.025)
Mean child age squared	-0.013 (0.001)	0.000 (0.001)	-0.004 (0.000)	-0.012 (0.002)
Middle school	0.267 (0.064)	0.072 (0.048)	0.194 (0.030)	0.208 (0.099)
High school	0.659 (0.126)	0.069 (0.085)	0.406 (0.053)	0.549 (0.186)
Year 1994	0.032 (0.023)	0.014 (0.019)	-0.002 (0.012)	0.030 (0.023)
Year 1995	0.139 (0.027)	0.024 (0.020)	-0.018 (0.012)	0.141 (0.027)
Year 1996	0.181 (0.026)	0.041 (0.020)	-0.026 (0.013)	0.182 (0.028)
Year 1997	0.248 (0.028)	0.020 (0.021)	-0.025 (0.013)	0.252 (0.029)
Year 1998	0.034 (0.033)	0.037 (0.023)	-0.009 (0.014)	0.032 (0.035)
First girl		3.486 (0.721)	-1.457 (0.453)	
× grandparents		0.107 (0.046)	0.009 (0.029)	
× mother Education		0.018 (0.005)	-0.015 (0.003)	
× mother age		-0.219 (0.042)	0.098 (0.027)	
× mother age squared		0.003 (0.001)	-0.001 (0.000)	
Constant	-5.115 (0.663)	-2.793 (0.529)	0.386 (0.332)	-4.919 (0.764)
R-squared	0.428	0.015	0.098	0.422
Partial R-squared		0.012	0.049	

Table 8 (continued)

	1	2	3	4
Regression stage	OLS	IV		
Dependent variable	ln <i>I</i>	First 1 D2	First 2 D3	Second ln <i>I</i>
<i>F</i> test		12.35	53.08	
(<i>p</i> value)		(<i>p</i> <0.01)	(<i>p</i> <0.01)	
Hansen <i>J</i> -statistic				1.93
(<i>p</i> value)				(<i>p</i> =0.59)


N=5,180. Robust standard errors, calculated by clustering for households, are displayed in parentheses.

The innovation of this paper is that I fully exploit differences in fertility (timing as well as number of children) due to the first child's sex. There are clearly different age profiles of fertility by the first child's sex. The age profiles implied by the first-stage estimates are depicted in Fig. 1. The first graph shows the marginal effects of *first girl* on the age profile of the probability of having exactly two children. The reference group is those with a son as the first child. The effects are graphed separately for three different education groups (middle-school graduates, high-school graduates, and college/university graduates). The age range for each educational group is chosen to guarantee that at least ten observations are available at each age to avoid any small-sample bias.

It is notable that the marginal effects are positive in the beginning and then decrease in age for all three educational groups. The positive effects are reasonable because having a girl as the first child would increase the likelihood of having a second child. Evaluated at the average age of the second childbirth for each group, the marginal effects are significantly positive. For middle-school graduates, the marginal effect is about 4 to 10% at the average age of 26.6, whereas it is 5 to 9% for high-school graduates at the average age of 27.6. The average age of second childbirth is highest (29–30) for college/university graduates. The marginal effect evaluated at the average age is, on the other hand, the lowest, 4 to 7%. Overall *first girl* increases the probability of having two children by 4–10%. The age profiles are consistent with the fact that higher-educated women delay childbearing, and they also prefer sons to daughters to a degree, but their preferences are weaker than those of lower-educated women.

The graphs also show that the marginal effect is decreasing and even becomes negative, which seems unreasonable. However, recall that we estimate the probability of having *exactly* two children. As a result, the marginal effect decreases as some of those who have a daughter at first fail to have a son again at the second attempt and try a third child. In particular, for middle-school and high-school graduates, the marginal effects are significantly negative after certain ages (29 for middle-school and 30 for high-school graduates). This suggests that, for these low-educated women, having a daughter at first increases the likelihood of having three children, compared to the reference group of households who have a son as their first child.

The bottom graph in Fig. 1 shows the marginal effects of *first girl* on the age profile of the probability of having three children based on the estimates in column 3. Even though the coefficient estimate for *first girl* is significantly positive, the graph shows that, combining it with the coefficient estimate for its interaction term

Fig. 1 Age profiles of fertility by the first child's sex and the mother's education 

with *mother age*, the marginal effects turn out to be consistently positive for all ages between 26 and 35 and actually increase with age. The profile for college/university graduates is not shown because there are few observations. This is not surprising because few college-educated women have three children (total 35 observations). Contrary to the previous graph, the marginal effects increase with age.¹³ The increasing profiles indicate that having a third child happens at a relatively later age. For middle-school graduates, *first girl* increases the probability of having three children by 9 to 11% at age 28.6, which is the average age when those women have a third child. The corresponding effect is 7 to 8% for high-school graduates at their average age 29.9. The results again confirm that higher-educated women postpone childbearing and have weaker preferences for sons.

Some results in column 4 of Table 8 are worth noting. First, investment in children's education is a normal good, and the income elasticity is quite inelastic, about 0.4. It confirms that children's education is a priority household expenditure. Those who live in metropolitan areas, especially in the capital Seoul, invest significantly more in children, although they have smaller number of children. Those in metropolitan areas invest more by about 5%, and those living in Seoul invest an additional 8 to 9%. This might reflect stronger preferences for child quality in these regions or more investment opportunities in cities, such as private tutoring and academic institutes. Not surprisingly, higher-educated parents tend to invest more in children's education. Investment is more responsive to the mother's education rather than to the father's education, which is consistent with the notion that child quality depends more on the mother-side characteristics. Higher-educated mothers have a smaller number of children, but they invest more in children's education. On the other hand, the father's education also increases investment, whereas it tends to increase the number of children.¹⁴

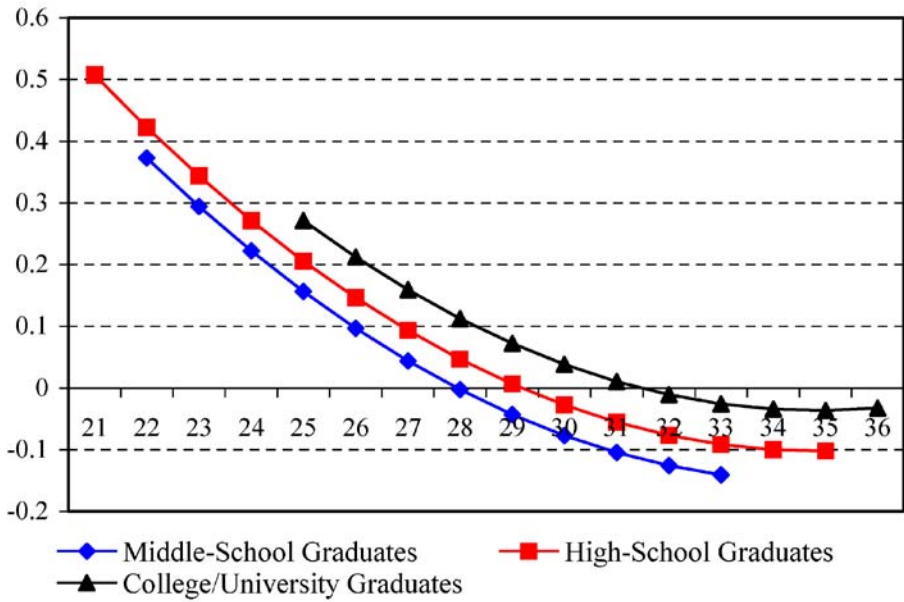
5 Conclusion

This study examines the effects of fertility on parental investment in children's education to test for the existence of a trade-off between child quantity and quality. The main contribution that this study makes is that it exploits exogenous variation in fertility due to son preferences. A consistent estimate of the causal effect of fertility on total investment in children's education is important for policy makers in developing countries. It is important to know the extent to which population policy

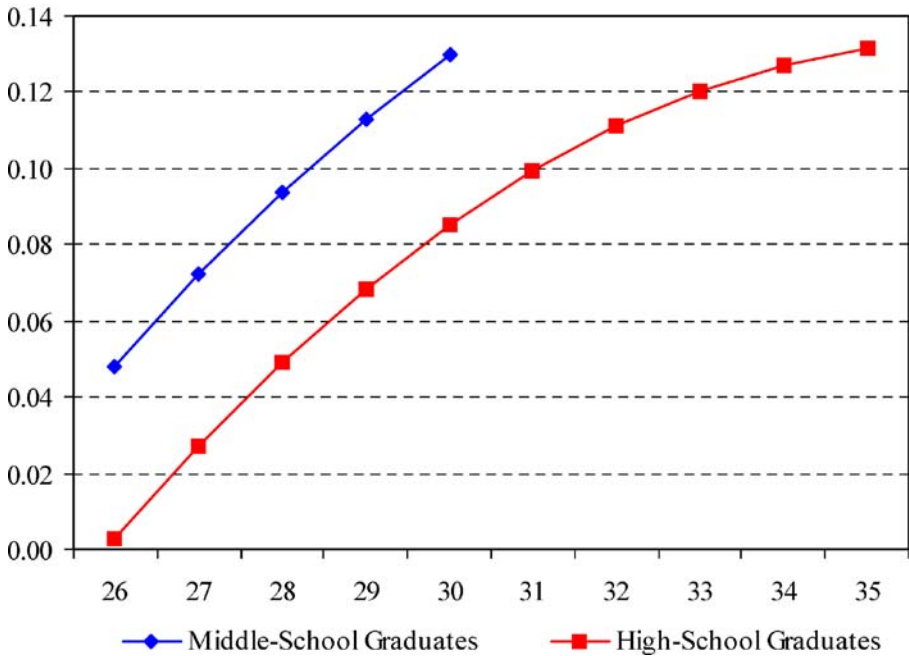
¹³ *Mother age* and *mother age squared* are insignificant in Table 6 where the dependent variable in the first-stage regression is a continuous measure of fertility (number of children). The two contrasting age profiles of the marginal effects are mixed in Table 6.

¹⁴ For robustness, I also experimented with a different instrument—the occurrence of twinning at the first birth. The validity condition in this study seems questionable, as the costs of educating children would differ for two singletons and two twins. Having two children at the same time may decrease per-child investment, in particular, when the household faces borrowing constraints. The estimate turns out to be significantly negative.

The Marginal Effects of First Daughter
on the Probability of Having Two Children over Mother's Age



The Marginal Effects of First Daughter
on the Probability of Having Three Children over Mother's Age



aiming at lowering fertility rate is effective in fostering human capital investment in the future labor force.

The results for South Korean households suggest that the observed trade-off exaggerates the true relationship due to unobserved heterogeneity. However, significantly, the trade-off still exists. In particular, as the number of children increases, the trade-off gets stronger. Slowing down population growth would be effective to increase per-child investment in education, particularly when fertility is high. If parents discriminated children against their sexes and invested more in sons, this conclusion would be strengthened.

The findings suggest that successful population policies in South Korea can explain in part why per-child investment in education has increased in the past decades. However, the magnitude of the causal effect of fertility on educational investment seems too small. The per-child investment in education in terms of the proportion of total household expenditure increased fourfold between 1981 and 2001. The fertility rate decreased from 2.6 to 1.3 for the same period. The IV estimates from the nonlinear model imply that, if the number of children per household decreased from 3 to 1, per-child investment should increase by 1.7 times. Thus, the decrease in fertility is not large enough to explain the fourfold increase in per-child investment over the last 20 years. There must have been other factors that decreased fertility and, at the same time, increased investment in education—for example, increasing parental concern about children's education or increases in the returns from human capital associated with technological changes.

Lastly, this study did not consider some important aspects in human capital investment for children, such as parents' time allocation and investment allocation among siblings. In particular, it would be interesting to examine both pecuniary and time investment in children in a comprehensive framework. I leave this important area for future research.

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