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ORIGINAL PAPER

Gender Gap in Health Status of Children in the Context of One-Child Policy in China: Is it Sibling Rivalry or Son Preference?

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Abstract This paper used data from 6 waves of the China Health and Nutrition Survey (CHNS) 1993–2009 to examine the effects of child gender, number of siblings, and sibling composition on children's health status, noting particularly any gender gaps, in the context of China's one-child policy. In the 1990s, the enforcement of the policy became less stringent, and ultrasound technologies, which enable prenatal gender selection, became more available. This led us to believe that the girls who were born were wanted by parents. In addition, growing income and industrialization should also narrow the gender gap in health status. Our results showed that on average, being a female decreased height-for-age approximately by 0.08 z-score as well as having an additional male sibling decreased height-for-age roughly by 0.17 z-score for a child under age 8. These results were particularly strong in the rural area but were non-existent in urban areas and among only-children. Our argument is that the gender gap was due to son preference and not sibling rivalry. Additionally, we found that having only one brother appeared to have a positive effect on girls' health status, compared to boys with another brother who tended to become rivals for household resources.

Keywords One-child policy · Health status · Gender gap · Sibling composition · Sibling rivalry · Son preference · Nutritional status · China

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Introduction

In order to curb the rapid population growth and combat poverty, China introduced the one-child policy in 1979. It has been regarded as one of the most controversial and aggressive family planning policies with far-reaching implications (Short and Zhai 1998) because of what the policy entailed: forced abortions and sterilizations, infanticide, sex-selective abortion, and infant abandonment. On the one hand, the policy was successful in reducing fertility for which it was designed and helped China achieve its unprecedented economic growth. On the other hand, the one-child policy engendered new problems as it transformed family size and social structure in China. Many scholars have raised an awareness of problems such as imbalanced sex ratio at birth, followed by looming marriage market failure and aging population, the consequences of which include social upheaval, prevalence of commercial sex activity and sexually transmitted disease, childless elderly (Ebenstein and Sharygin 2009) and "little emperors," spoiled only-children who did not grow up with many social interactions (Cameron et al. 2013).

Pro-male bias and male-biased sex ratios in China are well-documented and the prevalent practice of son preference is evident. Sen (1990) brought this issue to light with the estimate of 100 million missing women resulting from son preference. Short et al. (2001) explained that son preference in Chinese culture results from a patriarchal, Confucian view of gender. According to Confucian ideology, the ideal Chinese household consists of:

> as many generations of the male line as possible and as many male siblings as possible. In China's patrilineal society, only a son can carry on the family line. According to Confucian belief, one of the three grave

> unfilial acts is to fail to have a son, and in pre-revolutionary times, this was grounds for a man to divorce his wife. In addition, sons are considered a greater economic asset than daughters, who are usually lost to their natal household at marriage. Sons are expected to maintain financial and social ties with families throughout their lives and provide care for their aging parents. (Short et al. 2001, p. 915)

Li et al. (2011) estimated that the one-child policy resulted in about 7 extra boys per 100 girls for the 1991–2005 birth cohorts, which means the number of boys exceeded that of girls beyond the biologically stable range of gender ratio. Bulte et al. (2011) argued that son preference was the main driver in the reported male-biased sex ratio and the one-child policy was responsible for about half of it.

China's one-child policy was implemented in 1979 as a family planning policy. It was enforced at the provincial level. In principle, the one-child policy restricted urban couples to having only one child, while allowing additional children for some cases. Twins were an exception. Couples who were single children themselves and ethnic minorities were allowed to have a second child (Bulte et al. 2011). People who lived in remote areas could have a second child. Disabled couples and those engaged in hazardous occupations could also have a second child (Wu and Li 2012). The one-child policy was regulatory; however, the government employed incentives and disincentives to enhance the level of compliance. Those who complied with the policy were rewarded with extra food rations, better housing, health subsidies, and allotments of farmland (Bredenkamp 2009). Violation of the rule meant severe punishments. Violations required monetary penalties in the form of unaffordable fines and denied bonuses. For employees working in the public sectors, having an unsanctioned birth jeopardized employment status, opportunities for housing, and chances of promotion in the future (Wu and Li 2012). As a result of strict enforcement of the policy amidst a preference for sons and availability of sex-selection technology, what resulted was a male-biased sex ratio (Bulte et al. 2011; Ebenstein and Sharygin 2009; Li et al. 2011). In order to correct the imbalance in the sex ratio, in 2000, Chinese policymakers launched initiatives to subsidize parents who had only daughters, known as the Care for Girls campaign, and subsequently, in 2002, banned the use of ultrasound or other technologies for sex-selective abortion (Ebenstein and Sharygin 2009; Ebenstein 2011).

Many scholars have attempted to explore the impacts of one-child policy with limited success. There are three reasons for this. First, little data exist to pursue a "difference-in-difference" approach, i.e., estimate the sole exogenous impact of the policy on major outcome variables including anthropometric measurements by comparing periods before and after the policy. Second, as Short and Zhai (1998) concluded, no single one-child policy existed, i.e., the policy varied considerably from location to location. Local variability was expected even by the authority, who admitted in 1988 that the country was too vast and diverse for a uniform policy. Third, the stringency of the policy has varied considerably over time (Bredenkamp 2009). Enforcement was tightened for the first 4 years during which second births were effectively forbidden. By 1983, some coercive measures, mandatory IUD insertions, abortions, and sterilizations were reported. Fertility had been at about six births per woman up until 1970, but then had dropped to 2.2 births per woman by early 1980s (Bongaarts and Greenhalgh 1985). In early 1984, however, the severity was eased to suit local needs. In particular, there was a relaxation of rules for rural couples, who were allowed to have a second child if their first child was a daughter or disabled, implicitly re-enforcing societal preference for sons. Although the policy had been strictly enforced in urban areas, the implementation became less strict in the rural areas. Overall, due to the one-child policy, the fertility rate was greatly reduced as designed.

This paper examined the health status of Chinese boys and girls in the less stringent period of China's one-child policy. When assessing how girls have fared under the one-child policy, one can recognize two contradictory policy effects. On one hand, the one-child policy could have intensified the gender bias and hence, reduce investment in girls. Zhai and Gao (2010) noted that the one-child policy might aggravate discrimination against girls particularly in rural areas because parents perceive investing in girls' human capital to be risky since daughters are usually married out. On the other hand, the policy may have resulted in parents investing more in girls' health, because the girls who were born were most likely wanted by the parents (Short et al. 2001). Using data from the China Health and Nutrition Survey (CHNS) 1993–2009 waves, we examined the effects of child gender and number of siblings on children's health status, noting particularly any gender gaps, under China's one-child policy and amidst the tradition of son preference. We attempted to answer the following questions: Was there any difference in health status of only-child daughters and only-child sons? Was a one-child girl better off in health status than a girl with siblings, particularly a male sibling(s)? What sibling composition made a girl's health status worse off? If there was a gender gap, was this due to sibling rivalry for household resources or due to son preference?

Yu and Sarri (1997) examined health status of Chinese women from 1950s−1990s and found that infant mortality rates of females had been and remained high. Though female child mortality rates had declined dramatically, they still exhibited higher rates of deaths relative to boys, particularly in the rural areas. This was probably because daughters received less care and education, worked longer hours and were exposed to higher risk of morbidity and mortality. Song and Burgard (2008) studied children's growth trajectory in height over childhood and adolescence in China and particularly examining girls who survived infancy and early childhood found a male advantage in height over female counterparts in particular in the rural area during 1989–2000. Bian (1996) used data from 1990 and found number and structure of siblings were important determinants of parental monetary investment in their children in China. While parents in rural areas spent less on their children in absolute terms, they spent a higher proportion of their income compared to urban parents and all parents treated their sons and daughters differently. Zhai and Gao (2010) examined whether child gender and having siblings affected center-based care enrollment, which had increasingly become popular in China in recent decades due to the findings that center-based care tends to promote children's academic achievement compared to parental or grandparent care. This study found that only-children had higher chances of receiving center-based care than those with siblings. Though there was no evidence that child gender mattered, they found that the presence of older, school-age male siblings lowered preschoolers' chances of receiving center-based care. Lee (2012) examined the effects of one-child policy on gender equity in education. The author found that only children, particularly only-child girls, enjoyed greater opportunities for education, compared to children with siblings. There was no difference in educational attainment between only-child boys and only-child girls, while the gap between boys and girls both with siblings remained significant. Particularly, years of schooling for girls with brother(s) were 0.62 years lower than that of girls with sisters. Bredenkamp (2009) investigated what factors determined child nutritional status in the 1990s and examined the roles of the one-child policy and the health system reform. She found that being an only child increased nutritional status measured with height-for-age z-scores (*HAZs*) by 0.12 more than those of children with siblings. But, the nutritional outcome of only-children was not significantly better in higher income households, and there was no significant difference in nutritional status between only-child boys and only-child girls. Furthermore, access to quality healthcare was not significantly linked with better health status. Finally, using a March 2003 health survey and the National Fixed-Point Survey (NFS) from rural households in China, Gao and Yao (2006) examined the gender gap in access to health care. They found that girls under age 9 got significantly fewer treatments than did boys, and that girls' curative expenditures were sensitive to family income. Overall, these existing studies found that having more children, particularly male children, decreased resources available to family members and affected their human capital investments negatively. Only-children seemed to have enjoyed increased allocation of household resources relative to children with siblings, to the extent that the increased allocation may have adversely affected these only-children (Zhang et al. 2016).

Theoretically, we drew from human capital theory, the tradition of son preference, and theory of quantity-quality trade off. Human capital theory predicts that boys are favored by parents because the perceived returns to investing in sons are greater than investing in daughters in developing countries. On the other hand, son preference, which is a view of gender, is rooted in the tradition, culturally instilled and expressed in China. Son preference among many parents is clearly present, given the highly imbalanced sex ratio. It is a difficult task to determine what portion of gender gap in children's health status, if any, is attributed each to son preference or human capital investment. It is also often the case that in an emerging economy like China, rising income will give rise to increased human capital and narrower gender gaps (Garg and Morduch 1998). Moreover, Li et al. (2011) argued that the gender wage gap could explain only a small portion of the male-biased sex ratios in China. Therefore, should the gender gap in health status be found between sons and daughters, we would arguably attribute all of it to son preference.

The theory of quantity–quantity trade-off was formally developed by Gary S. Becker (1960). Quantity of children is negatively correlated with quality of children per family (Becker and Lewis 1973). Because parents must choose how to allocate available resources and time among their children, children often become rivals for binding household resources. Becker and Lewis (1973) explained that given the number of children per family:

> the cost of an additional child, holding their quality constant, is greater, the higher their quality is, because higher-quality children cost more. Similarly, the cost of a unit increase in quality, holding number constant, is greater, the greater the number of children, because the improvement in quality has to apply to more units. (p. 81–82)

This led to one of our hypothesis that when sibling size increases, average child quality measured in terms of health status should worsen for a given household budget. Our paper drew theoretical and empirical ideas from Garg and Morduch (1998), who investigated whether sibling gender composition and sibling rivalry determined child health outcomes in Ghana. They found that on average, children with all sisters (and no brothers) did roughly 25–40 % better on measured health indicators than children with all brothers (and no sisters). Garg and Morduch explored theoretical relationship between sibling composition and human capital. They speculated that there may be "spillover effects" for girls who have at least one brother, relative to girls with only sisters because having at least one son psychologically leads parents to treat daughters similarly to their sons. They also noted that:

> girls with only brothers may be treated differently from girls with at least one sister. Without sisters, a single daughter may be treated similarly to the boys in the family, but differences may widen once another girl is added to the family. (p. 476)

The authors called this "reference group effects." Whether there is a significant relationship between sibling composition and accumulated human capital is a case by case, empirical issue, and thus, to find the answer in the case of China is an interesting endeavor we hoped to accomplish in this paper.

Our hypothesis was that owing to the one-child policy, an only-child girl was better off in health status than a girl with siblings, particularly those with brothers, while the difference between only-child daughters and only-child sons was nonsignificant since one child was precious for parents regardless of gender. In particular, being one-child could be of great advantage, relative to other children with many siblings because one-child households would have more of household resources available for building up a child's human capital. Also, due to a strong attitude of son preference prevalent in China, we speculate that number of brothers had a larger negative impact on girls' nutritional status than number of sisters did. In the 1990s, the enforcement of the one-child policy became less stringent and ultrasound technologies, which enabled prenatal gender selection, became more available. This led us to believe that the girls who were born were "wanted" by parents (Short et al. 2001). Furthermore, economic and structural change towards industrialization coupled with increasing income and educational level often had a narrowing effect on gender gap (Garg and Morduch 1998; Li and Cooney 1993; Murphy et al. 2011).

In our study, we measured children's health status using objective measures such as anthropometric measures. Heights adjusted for age and gender had been used as an indicator of long term nutrition for children that could indicate long term health status of a child (Strauss and Thomas 1995). We found no significant difference in parents' health investment between only-child son and only-child daughter. However, we found parents' continued gender biases manifested in long-term health outcome for daughters under 8 years of age, particularly in the rural areas where parents had multiple offspring. We also found that girls with one brother, and no more than one brother, were likely to benefit from "spillover" effects, as opposed to boys who faced competition for household resources with another brother.

Our overall finding that girls were discriminated against in getting a fair share of household resources in rural areas is similar to the Gao and Yao (2006) paper. However, their study used a different outcome measure (access to health) and looked only at rural China. We used same anthropometric outcome as Bredenkamp (2009) paper. Yet, in contrast to Bredenkamp (2009) who used data from 1991 to 2000 and restricted data to children who were under the age of 12, our data were more recent covering 1993 through 2009. We also restricted our analysis to children who were under 8 years of age in order to minimize the risk of data reflecting genetic variation, as is the case of older children (Garg and Morduch 1998), and to avoid the risk of capturing indirect schooling and peer effects, making it an improvement on existing Chinese studies. Further, Bredenkamp investigated a difference in health outcomes between only-children and children with siblings but did not use number and composition of siblings to measure the gender gap as we did in this paper. Additionally to the extent of our knowledge, none of the previous studies explicitly addressed, as we did, whether the gender gap in health status under the one-child policy was due to sibling rivalry or son preference. Our paper also contributes to the existing literature by estimating the effects of number of siblings as well as composition of siblings on differential health status of girls and boys of age under 8 using most recently available data in the context of China's one-child policy.

The importance of investments in child health has been stressed by researchers such as Millimet and Wang (2011) and Strauss and Thomas (2007). Not only are adult health and skill development largely determined during the fetal and early childhood years, but adult stature is also positively associated with earnings at both the micro- and macro-economic levels. Hence, the results in our study could bode important implications for girls' health in later life, which might have far-reaching consequences for Chinese society (Osmani and Sen 2003).

Methods

Data

Data used in this study were obtained from the CHNS,[1] which is an international collaborative project between the University of North Carolina and the China National Institute of Nutrition and Food Safety. The CHNS employs a multistage, random cluster process to draw a sample of about 4400 households with a total of 26,000 individuals in seven provinces that are highly diversified in terms of social, economic, developmental, and demographic factors.

Our analysis was done using six waves of the CHNS survey from 1993, 1997, 2000, 2004, 2006 and 2009. Earliest waves did not have health related variables and the

[1] Available at http://www.cpc.unc.edu/projects/china.

Table 1 Variable definitions and descriptive statistics (sample restricted to age <8, 1993–2009)

Variable	Definition	Descriptive statistics				
		Mean	SD	Min	Max	N
HAZ	Height (cm) for age (months)—z-score calculated using *zanthro*	−0.90	1.27	−4.99	4.04	3077
Stunted	HFA being 2 standard deviations below the reference population	0.14	0.35	0.00	1.00	4170
Fatherheight	Biological father's height(cm)	167.06	6.33	145.20	186.00	3788
Motherheight	Biological mother's height(cm)	156.42	6.13	75.00	176.10	3940
Fatheredu	Father's cumulative years of schooling	8.84	3.10	0.00	20.00	3825
Motheredu	Mother's cumulative years of schooling	7.61	3.61	0.00	20.00	3693
Male_sibN	Number of male siblings	0.29	0.50	0.00	4.00	4170
Female_sibN	Number of female siblings	0.38	0.64	0.00	4.00	4170
Birth order	A child's rank by age among his/her siblings	1.49	0.69	1.00	5.00	4170
lhhinc_cpi	Log of household monthly income inflated to 2009	9.50	1.01	1.83	13.61	4122
Urban	Dummy variable: 1 for urban; 0 otherwise	0.27	0.44	0.00	1.00	4170
Urban female	Dummy variable: 1 for urban & female; 0 otherwise	0.2859	0.4519	0	1	1875
One child	Dummy variable: 1 for one child; 0 otherwise	0.49	0.50	0.00	1.00	4170
1997	Dummy variable: 1 for 1997 survey; 0 otherwise	0.22	0.41	0.00	1.00	4170
2000	Dummy variable: 1 for 2000 survey; 0 otherwise	0.12	0.32	0.00	1.00	4170
2004	Dummy variable: 1 for 2004 survey; 0 otherwise	0.14	0.34	0.00	1.00	4170
2006	Dummy variable: 1 for 2006 survey; 0 otherwise	0.12	0.32	0.00	1.00	4170
2009	Dummy variable: 1 for 2009 survey; 0 otherwise	0.13	0.34	0.00	1.00	4170
Age	Age in month	55.60	27.40	0.13	97.33	4170
Age2	Age in month squared	3842.51	2903.22	0.02	9473.78	4170
Female	Dummy variable: 1 for girls; 0 otherwise	0.45	0.50	0.00	1.00	4170
EarnerN	Number of income earners in the household	2.28	1.16	0.00	8.00	4170
Sample Size (N)	Number of observations for age <8, 1993–2009					4170

latest available 2011 round was not publicly available at the time we did this analysis. Our dataset consisted of a sample size of 12,308 children under 18 years of age, of which the sample of children under 8 years of age was 4170. We analyzed health status using height for age measures. Height-for-age (HFA) is a measure of long-term health status, and we used *HAZ*, adjusted for gender and age in months, to reflect HFA. We used *zanthro* in Stata 11 to come up with the *HAZ*. These scores were calculated based on WHO/UK/CDC growth charts (Vidmar et al. 2004). Following most studies, we used the WHO[2] criterion, where "stunting," the indicator of extreme health outcomes, is defined as having a *HAZ* more than two standard deviations (SDs) below the median, (i.e., z-score less than −2) of the international reference. We excluded children with *HAZ*s that are above 3.0 and below −5.0 from the analysis on the ground that those values are implausible (Bredenkamp 2009; WHO 1995). Our sample was restricted to children under 8 years of age to examine gender gap in long term health status by minimizing other confounding effects due to growth spurts, genetic, and peer impacts. Due to sample size issues, to analyze sibling composition effects, we used the whole sample of children below 18 years of age. Table 1 provides the definition of the variables used in our analysis. Of all children under eight in the sample, 14.2 % were stunted, 26.7 % resided in the urban area, 45 % were female, and 48.8 % were only-children.

Descriptive analysis of the sample revealed gender bias, which was further analyzed by empirical models later in the study. Table 2 presents the percentage of stunting by gender among only children and children with siblings for the sample restricted to age under 8. The first thing to note is that our data confirmed the imbalanced sex ratio widely discussed by many papers. Of children who were under 8 years old, the sex ratio of male to female deviated far from the biologically stable range of 103 to 107 (Li et al. 2011) for each year of the survey. The ratio had even increased from 122.33 in 1993 to 134.50 in 2009, despite the government's effort through the legislative initiatives as discussed earlier in the introduction. While these numbers seem high, they were not implausible as these numbers were from a survey sample restricted to below 8 years of age. Other studies found similar sex ratios for survey data from China (Guilmoto 2012; Zhu et al. 2009). Second, the number of only-children had

[2] See The World Health Organization http://www.who.int/en/.

Table 2 Percentage of stunting by gender among only children and children with siblings (sample restricted to age <8, 1993–2009)

	Male						Female						All					
	1993	1997	2000	2004	2006	2009	1993	1997	2000	2004	2006	2009	1993	1997	2000	2004	2006	2009
Number of one child	203	225	166	183	164	168	150	214	140	159	136	127	353	439	306	342	300	295
% of one child	31.67	46.01	63.12	57.37	59.64	54.55	28.6	50	62.5	63.1	62.4	55.46	30.3	47.87	62.83	59.89	60.85	54.93
% of stunting	10.84	6.22	10.24	3.28	12.8*	4.76	12	7.48	7.14	8.81	5.15	3.15	11.33	6.83	8.82	5.85	9.33	4.07
Number of child w/siblings	438	264	97	136	111	140	374	214	84	93	82	102	812	478	181	229	193	242
% of child w/siblings	68.33	53.99	36.88	42.63	40.36	45.45	71.4	50	37.5	36.9	37.6	44.54	69.7	52.13	37.17	40.11	39.15	45.07
% of stunting	25.34	20.45	22.68	13.97	12.61	2.86	25.9	21.5	26.2	24.73	17.1	8.82	25.62	20.92	24.31	18.34	14.51	5.37
Number of stunted children	133	68	39	25	35	12	115	62	32	37	21	13	248	130	71	62	56	25
% of stunted children	20.75	13.91	14.83	7.84	12.73	3.9	22	14.5	14.3	14.68	9.63	5.68	21.29	14.18	14.58	10.86	11.36	4.66
Total	641	489	263	319	275	308	524	428	224	252	218	229	1165	917	487	571	493	537
Sex ratio (male to female)													122.3	114.3	117.41	126.6	126.15	134.50
Ave. HH income (yuan)													12,040	14,594	18,349	25,479	30,349	42,991

An only child is defined as someone under the age of 18 without siblings who are younger than 18

*The percentage of stunting among male only children in 2006 is 12.8 %, which seems peculiarly high. We suspect this might be due to measurement errors

increased dramatically. In 2009, 54.93 % of children were only-children compared to 30.3 % of children in 1993. Lastly, the percentage of stunting among children under eight fell from 21.29 % in 1993 to 4.66 % in 2009, by 16.63 percentage points, as the percentage of only-children and the average household income both increased. The comparison of percentage of stunting between only-children and children with siblings revealed that in 1993, the percentage of stunting for children with siblings was much higher than that of only-children, 25.62 vs. 11.33 %; however, by 2009, the gap had narrowed to a 1.3 percentage point difference. Nonetheless, taking a closer look at the gender difference between boys and girls unveiled an interesting story. Among only-children, the gender difference in stunting was minimal. What was interesting was that, among children with siblings, the percentage of stunting among girls was persistently higher than that of boys throughout the years; besides, although the percentage of stunting between boys and girls differed only by 0.6 percentage points in 1993, the gap had widened as large as 5.96 percentage points by 2009 (2.86 % for male children vs. 8.82 % for female children). This aberrant divergence is a potential manifestation of differences in the treatment of sons vs. daughters, and gave us a strong ground to investigate whether female children were disadvantaged due to either sibling rivalry and/or son preference, even after they were born.

Table 3 provides the average of *HAZ* for four groups of children under age 8: only-child boy, only-child girl, boy with sibling(s), and girl with sibling(s). Apparently, being only-children gave an advantage in nutritional status in the absence of competition. HAZ was very similar among only children regardless of gender; however, the mean height for girls with sibling(s) was lower than for boys with sibling(s) by z-score 0.18. Although the result of the mean test suggested that the mean HAZ between boys and girls among children with sibling(s) were not significantly different, the difference of z-score 0.18 warranted a further investigation for differential treatment for boys and girls by parents.

Table 3 Average of height-for-age z-scores for four groups of children under 8 years of age

Mean HAZ	One-child			Child w/sibling(s)		
	Boy	Girl	Test of equality (p value)	Boy	Girl	Test of equality (p value)
	−0.53	−0.58	(0.43)	−1.14	−1.31	(0.22)

Variance ratio test, which preceeded two sample *t* test, is based on 10 % significance level

Empirical Analysis

Main Model (OLS)

For the estimation of gender gap in health status under the one-child policy, the main model we used is a set of pooled ordinary least square (OLS) regression, clustered at the individual level. The model estimated is:

$$\begin{aligned} HAZ_{it} = {} & \beta_0 + \beta_1 \cdot female_{it} + \beta_2 \cdot male\_sibN_{it} \\ & + \beta_3 \cdot female\_sibN_{it} + \beta_4 \cdot fatherheight_{it} \\ & + \beta_5 \cdot motherheight_{it} + \beta_6 \cdot motheredu_{it} \\ & + \beta_7 \cdot lhhinc\_cpi_{it} + \beta_8 \cdot age_{it} + \beta_9 \cdot age2_{it} \\ & + \beta_{10} \cdot onechild_{it} + \beta_{11} \cdot urban_{it} + \beta_{12} \cdot 1997_i \\ & + \beta_{13} \cdot 2000_i + \beta_{14} \cdot 2004_i + \beta_{15} \cdot 2006_i \\ & + \beta_{16} \cdot 2009_i + \sum_{i=1}^{8} province_{it} + \varepsilon_{it} \end{aligned}$$

where i denotes individuals and t denotes the survey year. The term, ε_{it}, represents i.i.d. random error. We call this the pooled OLS. The dependent variable is *HAZ*. The main variables of interest are *female* and *male_sibN*. The binary variable *female* is to estimate the effect of being a female on her health status, relative to a male child. We expected the coefficient of female to be negative because of the pro-male bias. The variables of number of brothers and sisters (*male_sibN* and *female_sibN*) were constructed to measure the impacts of having additional sibling on child's health status. Having more brothers (*male_sibN*) was expected to impose a larger negative effect on child's health than having more sisters (*female_sibN*).

Other explanatory variables in the regression are as follows. *Motherheight* and *fatherheight* were included to control for genetic endowments of their children. For this purpose, children in the sample were matched with their biological parents. As many past studies had emphasized, parents' years of education in particular mother's education (*motheredu*), was expected to play a positive role in children's health. The dummy variable, *urban*, was designed to control regional impacts on health. Other control variables are: *lhhinc_cpi*[3] (in 2009 yuan) and the dummy variable, *onechild*. Both were expected to have a positive effect because household income increases resources available for a given family, and only-child does not face any competitions for household resources. Age measured in months (*age*) and age squared (*age2*) were included to capture the growth-faltering of children commonly detected in developing countries (Bredenkamp 2009). Five survey years[4] (*1997, 2000, 2004, 2006, 2009*) were included to control the economic trends of the years and the characteristic of how each survey was done. We treated *1993* as the reference year. Province dummies were used to control for province-level heterogeneity. We treated province 9 as the reference. Our estimations included the heteroskedasticity-robust standard errors, because the estimator of pooled OLS produces incorrect standard errors and typically, overstates the reliability of the estimator[5] (Hill et al. 2010). This specification also gave the best R^2 (=0.2799) among alternatives.

Sensitivity Analysis

We introduced various models to study the sensitivity of our estimates. The fixed effects (FE) and random effects (RE) models were applied in order to control for province-specific unobserved heterogeneity. There are a number of reasons for unobserved province-specific components to affect child's height-for-age. Different provinces may have different attitudes towards gender. Different provinces may have different anticipations for gender roles for economic, geographic and cultural reasons. Different provinces may probably have different types or degrees of community services which nourish local children. Also, the provinces differed as to how stringent the enforcement of one-child policy was. These factors usually change very slowly over time, if any at all, and hence, the FE and RE models can eliminate these unobserved effects specific to the provinces and produce consistent estimates of gender bias, if any.

The income variable, *lhhinc_cpi*, is potentially endogeneous because of the simultaneous determination of household income and child's health status. We emulated Bredenkamp (2009) and employed the number of income earner (*earnerN*) as an instrument variable (IV) to estimate a Two Stage Least Square (2SLS) model, though doing so undermines the efficiency of the OLS estimates for the sake of gaining the consistency of the IV estimates. Despite different model specification, the instrument, *earnerN*, was very strong with a t-statistic $= 11.58 > 3.3$ and an F value $= 134.21 > 10$ in the first stage regression and hence, a reliable instrument. Number of income earner, though correlated with income, should not directly impact nutritional status of children. Bredenkamp using data from the same time period used an over-identification test to confirm the instrument's exogeneity and thus, validity.

Further, the number and nutritional status of children may be simultaneously determined and hence, the variables, *onechild, male_sibN*, and *female_sibN* in our model may be endogenous. Although under one-child policy, couples could

3 Real household monthly income is a natural log form because it gave a better fit in terms of R^2.

4 The number of observations used in the regression models is 930, 438, 279, 325, 264, and 297 for the years 1993, 1997, 2000, 2004, 2006, and 2009 respectively. .

5 We carried out Breusch-Pegan test and White test for heteroskedasticity on the pooled OLS model. Both the tests confirmed the presence of heteroskedasticity.

Table 4 Proportion of only children vis-à-vis children with siblings for the nine provinces, 1993–2009

	Liaoning	Heilongjiang	Jiangsu	Shandong	Henan	Hubei	Hunan	Guangxi	Guizhou
One-child	0.664	0.71	0.664	0.522	0.390	0.374	0.437	0.315	0.341
Child w/siblings	0.336	0.29	0.336	0.478	0.610	0.626	0.563	0.685	0.659
Observations	903	1158	1185	1145	1585	1496	1322	1820	1694

have a limited number of children imposed by the rules to which they were subjected,[6] there was temporal (Table 2) as well as spatial variations (Table 4) in the implementation of one-child policy. The number of children therefore was most likely a subject to parental choice, hence an endogenous variable. However, the variation in the number of children was most likely influenced by household budget constraints (ability to pay fines for additional children). Thus correcting for endogeneity around household income as described above helps reduce any significant endogeneity bias. Also, under the coercive one-child policy, it is not entirely impossible to assume that the number of children was exogenous in the present study.[7]

Subgroup Analysis

In order to further investigate how gender gap played out for specific groups of children, the OLS model was further applied to a sub-group analysis, namely: urban, rural, only-child, child with siblings, boys, girls, only-child living in the urban area, only-child living in the rural area, child with siblings in the urban area, and child with siblings in the rural area.

Impacts of Sibling Composition

Lastly, we examined how the gender gap manifested among children with specific sibling composition. The sample of all children under 18 years of age was used in order to reasonably estimate as many groups as possible. The 25 dummy sibling composition groups were constructed as follows: a group of children with:

- no brothers and no sisters, 1 brother and no sisters, 2 brothers and no sisters, 3 brothers and no sisters, 4 brothers and no sisters
- no brothers and 1 sister, 1 brother and 1 sister, 2 brothers and 1 sister, 3 brothers and 1 sister, 4 brothers and 1 sister
- no brothers and 2 sisters, 1 brother and 2 sisters, 2 brothers and 2 sisters, 3 brothers and 2 sisters, 4 brothers and 2 sisters
- no brothers and 3 sisters, 1 brother and 3 sisters, 2 brothers and 3 sisters, 3 brothers and 3 sisters, 4 brothers and 3 sisters
- no brothers and 4 sisters, 1 brother and 4 sisters, 2 brothers and 4 sisters, 3 brothers and 4 sisters, 4 brothers and 4 sisters

Of these groups, six sibling composition groups had sufficient number of observations for a robust estimation. The specification differs slightly from the main OLS as the variables related to the number of siblings, namely, *male_sibN, female_sibN*, and *onechild*, were dropped to evade perfect multi-collinearity.

Results

Main Results: OLS with Sensitivity Analysis

Table 5 presents the results for the main pooled OLS model, along with the results for FE, RE, and the 2SLS model.[8] The coefficients of the dummy variable indicating a female child (*female*) were consistently negative, and significant in the OLS model and 2SLS, but not in the FE and RE model. Yet, the magnitudes of the coefficient were comparable and consistent across all the models. This means that holding all else constant, being female was negatively correlated with *HAZ* with female children's scored lower than male children's by 0.08. While the coefficients of number of female siblings (*female_sibN*) were small and nonsignificant, coefficients of number of

[6] Although ethnic minority is an exception to the one-child policy, minorities comprised very small portion of our sample and thus, exclusion of minorities did not change the estimates.

[7] We tried many variables for the potential instrument for *onechild*. We list a few of them: "The first child is a girl," "Parents have a sibling," "Mother has a career," "The number of hours mother works," "The number of siblings parents have," "The years of marriage." Nevertheless, none produced consistent results, signifying that number of children in China may be exogenous under the influence of the one-child policy. Furthermore, Lee (2012) concluded that a potential problem of variable *onechild* is not a serious issue as she found that the IV estimate and OLS estimate of *onechild* were very similar, which also favors our speculation that the potential endogeneity problem between children's health status and fertility may be minimal.

[8] The Variance Inflation Factor (VIF) was used to assess the degree of multicollinearity. Using the pooled OLS model, the VIF values for all the variables but *age* and *age2* were much $^{<}$10 and the mean VIF was 6.77, which did not exceed 8, and therefore, multicollinearity was unlikely to be an issue.

Table 5 Gender gap in health status; main OLS results in comparison with FE, RE, and 2SLS results (sample restricted to age <8, 1993–2009; dependent variable: HAZ)

Independent variable	Pooled OLS	FE	RE	2SLS
Female	−0.081†	−0.081	−0.076	−0.080†
	(0.089)	(0.184)	(0.129)	(0.09)
Number of male siblings	−0.166*	−0.166*	−0.183*	−0.163*
	(0.015)	(0.047)	(0.021)	(0.016)
Number of female siblings	−0.037	−0.037	−0.045	−0.036
	(0.441)	(0.534)	(0.451)	(0.477)
Mother's height	0.035***	0.035**	0.044***	0.036***
	(0.000)	(0.002)	(0.000)	(0.000)
Father's height	0.046***	0.046***	0.052***	0.046***
	(0.000)	(0.000)	(0.000)	(0.000)
Mother's education	0.018*	0.018*	0.016*	0.019*
	(0.013)	(0.018)	(0.012)	(0.04)
Log of household income	0.03	0.03	0.028	0.013
	(0.256)	(0.254)	(0.325)	(0.913)
Urban	0.155**	0.155**	0.135***	0.159*
	(0.003)	(0.001)	(0.000)	(0.011)
One child	0.109	0.109	0.152	0.115
	(0.171)	(0.293)	(0.134)	(0.148)
Age	0.022**	0.022**	0.022***	0.022**
	(0.001)	(0.007)	(0.000)	(0.001)
Age squared	−0.0002***	−0.0002**	−0.0002***	−0.0002***
	(0.000)	(0.004)	(0.000)	(0.000)
R^2	0.2916	0.2042	0.2028	0.2915
Sample size	2516	2516	2516	2516

Each equation includes an intercept term, eight province dummies, and 5 year dummies. Numbers in parentheses are p values associated with heteroscedasticity-robust standard errors clustered at the individual level

Statistical significance level: †$p<.1$; *$p<.05$, **$p<.01$, ***$p<.001$

male siblings (*male_sibN*) were negative and highly significant in all models. Furthermore, the magnitudes of the negative effect of having one more brother were much larger than the negative effects of having one more sister. The effect size differed between the RE and the rest of the models. Regression outputs of the FE and RE models report the estimated values of the error serial correlation. These were 0.023 for FE and 0.000 for RE indicating that the province-specific unobserved heterogeneity elucidated a small portion of the total variation in the residuals (Wooldridge 2010). Hence, serial correlation was of little concern, and the results from the pooled model could be considered reasonably reliable. Nonetheless, because the FE model controls for the unobserved, province-specific heterogeneity, and some of the explanatory variables are likely to be correlated with province FE, the magnitude of the coefficient of number of male siblings (*male_sibN*) in the FE model is a closer estimate to the true value. Thus we could conclude that all else equal, an additional brother in the household would decrease a child's *HAZ* by as much as 0.17.

In addition, the coefficients of parents' height were positive and significant as expected. As noted by numerous studies, the coefficients of mother's education were persistently positive and significant, suggesting mother's education to be an important determinant for a child's health. The sign and significance of coefficients of age and square of age confirmed that on average, the height for children under eight in China exhibited growth-faltering commonly observed in developing countries. The coefficient of *urban* was positive and significant, meaning that children who resided in the urban area had better health status than their counterparts in rural areas. Results also indicated that there were positive benefits to being one child. Though nonsignificant, the size effect of the coefficient of being only-child (*onechild*) was z-score 0.11 in all the models except the RE model. This value is very close to the estimate that Bredenkamp (2009) reported, which is z-score 0.119. The coefficient of log of household income (*lhhinc_cpi*) was positive but not significant, suggesting that household income was not a strong determinant of a child's nutritional status on average beyond the overall upward economic trends.

Table 6 Pooled OLS regression results for each sub-group (sample restricted to age <8, 1993–2009; dependent variable: HAZ)

Independent variable	Urban	Rural	One-child	W/siblings	Boy only	Girl only
Female	−0.038 (0.671)	−0.104† (0.063)	−0.067 (0.352)	−0.095 (0.131)		
Number of male siblings	−0.116 (0.381)	−0.159* (0.038)		−0.158* (0.027)	−0.184* (0.026)	−0.148 (0.202)
Number of female siblings	0.056 (0.648)	−0.048 (0.352)		−0.033 (0.505)	0.012 (0.848)	−0.093 (0.200)
Mother's height	0.046*** (0.000)	0.031*** (0.000)	0.029*** (0.000)	0.039*** (0.000)	0.040*** (0.000)	0.030*** (0.000)
Father's height	0.045*** (0.000)	0.046*** (0.000)	0.041*** (0.000)	0.050*** (0.000)	0.041*** (0.000)	0.051*** (0.000)
Mother's education	0.031* (0.040)	0.013 (0.117)	0.041**	0.002 (0.843)	0.022* (0.016)	0.014 (0.250)
Log of household income	−0.043 (0.410)	0.050 (0.103)	0.021 (0.617)	0.030 (0.341)	−0.019 (0.558)	**0.095*** (0.024)
Urban			0.069 (0.380)	0.249*** (0.000)	0.148* (0.028)	0.158† (0.062)
One child	0.076 (0.639)	0.146 (0.111)			0.099 (0.323)	0.135 (0.319)
Age	0.003 (0.838)	0.027*** (0.000)	0.015 (0.142)	0.029** (0.001)	0.015† (0.093)	0.031** (0.002)
Age squared	−0.00004 (0.732)	−0.0002*** (0.000)	−0.0001† (0.088)	−0.0002** (0.001)	−0.0001† (0.077)	−0.0003** (0.001)
R^2	0.3604	0.2655	0.1921	0.2811	0.3113	0.2873
Sample size	662	1854	1107	1409	1405	1111

Each equation also includes an intercept term, eight province dummies, and 5 year dummies. Numbers in parentheses are p values associated with heteroscedasticity-robust standard errors clustered at the individual level

Statistical significance level: $^{\dagger}p<.1$; $*p<.05$, $**p<.01$, $***p<.001$

As discussed before, number of children and household income may be endogenous. Although household income turned out to be not a significant contributor to child's height-for-age after the year FE were controlled, we estimated a 2SLS model, using an instrument, number of income earners in the household (*earner*) for household income. We used number of earners as an instrument based on the validity test reported in Bredenkamp (2009) that used data from same period (of 1990s) as in this study. We expected that number of income earners in the household would positively impact nutritional status of children but only through the availability of higher incomes earned that would allow households more resources or access to high quality day care. In multimember households such as in the Chinese context, there was no reason to assume that number of income earners critically compromised sufficient number of caregivers for children hence possibly health of children. In the 2SLS model, the size of coefficient of household income, 0.013, was smaller than those in the other models, which was about 0.03, implying that household income was in fact endogenous. However, we stress here that the 2SLS model was estimated to check the sensitivity of the estimates obtained by the main OLS model. The important result from the 2SLS model is the consistent size of the estimates of the variables obtained by the main OLS model even after part of the endogeneity problems were accounted for. In particular, the coefficients for female and number of male siblings remained consistent and robust across all specifications.

Sub-Group Analysis

The OLS estimates for sub-groups of children are shown in Tables 6 and 7. According to Table 6, even though the coefficients for female and number of male siblings for children living in urban areas (Column 1) and only-children (Column 3) were negative, they were statistically nonsignificant. In contrast, according to Column 2 in Table 6, on average, being female in the rural area reduced HAZ by 0.104. Also, for both boys only and girls only samples, having an additional male sibling had a negative impact on *HAZ*s but the result was statistically significant for the girls only sample (Column 6). Further, household income had a positive impact on girls' *HAZ* for the girl only sample, as shown in Table 6, Column 6. According to results disaggregated by rural/urban and with/without siblings as reported in Table 7, Column 4, the negative effect of having one more brother in rural households on *HAZ* was 0.158, and additionally, being female negatively affected a girl's *HAZ* by approximately 0.132 relative to male children in rural areas. This gender gap in rural areas was about 0.051 (=0.132 − 0.081) larger in magnitude than the result of the pooled OLS for the full sample seen in Table 5, Column 1. Average *HAZ* for children with siblings in rural areas is 1.24. So 0.132 lowerHAZ meant 10.6 % lower height for these girls compared to boys.

Impacts of Sibling Composition

As explained earlier, 25 groups were constructed as presented in Table 8 in order to estimate the effect of being

Table 7 Gender gap for rural and urban samples; pooled OLS regression results (sample restricted to age <8, 1993–2009; dependent variable: HAZ)

Independent Variable	Urban& Onechild	Urban & w/siblings	Rural & Onechild	Rural & w/siblings
Female	−0.087 (0.469)	−0.008 (0.954)	−0.060 (0.500)	−0.132† (0.066)
Number of male siblings		−0.146 (0.341)		−0.158* (0.045)
Number of female siblings		0.020 (0.886)		−0.051 (0.342)
Mother's height	0.041** (0.002)	0.058*** (0.000)	0.024* (0.018)	0.034*** (0.000)
Father's height	0.042** (0.001)	0.042*** (0.000)	0.038*** (0.000)	0.051*** (0.000)
Mother's education	0.037† (0.082)	0.029 (0.144)	0.045*** (0.002)	−0.003 (0.792)
Log of household income	−0.020 (0.787)	−0.100 (0.115)	0.023 (0.668)	0.060 (0.106)
Age	−0.012 (0.534)	0.029 (0.118)	0.026* (0.027)	0.030** (0.003)
Age squared	0.0001 (0.582)	−0.0002† (0.086)	−0.0002* (0.014)	−0.0002** (0.002)
R^2	0.2489	0.4425	0.1687	0.2571
sample size	387	275	720	1134

Each equation also includes an intercept term, eight province dummies and 5 year dummies. Numbers in parentheses are p values associated with heteroscedasticity-robust standard errors clustered at the individual level

Statistical significance level: $^{\dagger}p < .1$; $*p < .05$, $**p < .01$, $***p < .001$

female on height-for-age relative to being male using the OLS model among children with specific sibling compositions. Yet, only 6 groups out of 25 groups that had adequate number of observations were examined. For instance, a dummy variable of the group of 618 observations in Table 8 takes 1 for a child with 1 brother and 1 sister, and 0 otherwise. Table 8 presents these results. It can be seen that without any sisters, girls who had just one brother was better off by HAZ 0.096 than boys who had one brother. Even though statistically not significant, the signs of the other coefficients were all in line with our expectations. A girl with no brothers and two sisters was worse off than a boy with the same sibling composition by 0.131 z-score. We surmise that this was because parents without a male offspring in a multi-children household lacked the willingness to invest in their female children's health.

Discussion

Our results confirmed our hypothesis that there was significant gender gap in health status in rural China with girl children worse off than boys and having a male sibling reinforced this negative effect. The fact that having sisters had no significant effect but having brothers did points to the reason being son preference rather than sibling rivalry. If this were a sibling rivalry effect, it would not have mattered whether it was male or female siblings competing for resources. However, since the male effect was so strong, this was indication of son preference where brothers took away from health investments in girls, creating a strong negative impact on girl's long term health status. Children who resided in urban areas had better nutritional status than their counterparts in rural areas.

In China where one child policy was strictly enforced, the expectation would then be that children who were only boys or only girls were wanted by their parents regardless of gender, hence we should not expect to see a gender gap in health investments and thus in long term health status. This was indeed the case as our results showed no significant gender gap in health status among single children. The anthropometric outcomes did not differ significantly by gender in urban areas as well, while they did among children in the rural areas and children with siblings. These results for only-children and children living in the urban area were, however, expected since it is natural to think that parents invest in their only-child's health regardless of gender. Also, the results for the urban area were as expected for communities where the majority have only-children and are relatively wealthier as well as the environment which makes it difficult for parents to make a conspicuous difference in their children's health investments by gender. In stark contrast, girls who lived in the rural area and had siblings were clearly worse off. Was this difference between boys and girls because of sibling rivalry or because of son preference? The fact that we did not see a significant impact of having a female sibling on the health of children but we saw a strong negative impact of a male sibling seems to suggest that the discrimination happened not because of sibling rivalry of resources but because of a strong male preference amongst parents.

Our results also show that having another male sibling had a peculiarly high negative impact on boys' health status. We speculate that this could be a result of parental preference

Table 8 Impact of sibling composition on female height-for-age (sample restricted to age <18, 1993–2009)

Number of sisters	Number of brothers			
	0	1	2	3
0	0.03	0.096†	0.089	–
	(3945)	(1970)	(186)	(7)
1	−0.003	−0.036	–	–
	(1688)	(618)	(46)	(4)
2	−0.131	–	–	–
	(295)	(91)	(9)	
3	–	–	–	–
	(28)	(20)		
4	–	–	–	–
	(15)			

For each cell, the coefficient of "female" is reported and the number in parentheses is the sample size

Statistical significance level †$p<.1$; *$p<.05$, **$p<.01$, ***$p<.001$

(–) Signifies that the model was not estimated for the groups due to insufficient sample size

for particular birth orders but found no such evidence.[9] This effect could also be a result of other factors. For example, under one child policy, if the first born was a boy, in general, a family was discouraged from having a second child unless the family was from a minority ethnic group, or if the first son had a serious disease or disability. Having the second or third son would incur hefty fines. The observed effect could be associated with the minority ethnic group status (many of this group live in hard-to-reach rural areas), chronic disease of the first born, and even the fine. Minority Chinese families may be in worse health to begin with and the effect of having a second son may be an over-estimate of the negative impact. As mentioned earlier in the paper, we did exclude ethnic minorities to check for this effect but the results did not change possibly because the sample size of ethnic minorities is too small. Also, if the first son had a chronic disease or if the family had to pay a hefty fine for a second child, these could be reasons for impact of a second son to show up as a negative impact on health of the first son (because the first son was of worse health to begin with or families had less resources), rather than the impact being truly due to the family's preferential treatment of the son of higher birth order. Result obtained in this analysis could also be combined effect of all of these factors together.

Also, 28.59% of female children lived in the urban area and 71.41% lived in the rural areas (Table 1). The strong gender gap results, particularly in the rural areas, clearly illustrate that girls were of secondary importance to parents in Chinese households in rural areas where a strong penchant for son is generally expressed. Interestingly, household income had a positive impact on girls' HAZ in the girl-only sample. For example, using our sample, for a family with an average income, a 10% increase in household monthly income (= approximately 2012 yuan, which is about US$295 in 2009 dollars) was estimated to increase a girl's HAZ by approximately 0.01. This empirical result is consistent with the economic assumption that "parents' aversion to the unequal treatment of their children increases with income" (Garg and Morduch 1998, p. 473). Other factors likely influencing a girl's HAZ, such as urban residence or composition of siblings, are difficult to change in a short period of time by public policies, given that son preference seems tenaciously instilled in rural areas. The implication of these results is that, in the presence of son preference, improving a girl's nutritional status could only be successful by targeting and expanding household budgets.

Finally, we also examined gender gap in health status by sibling composition and conclude that girls who had just one brother were better off than boys who had one brother. This result is consistent with what Garg and Morduch (1998) called "spillover effects," that is, there may be spillovers for girls who have at least one brother, relative to girls with only sisters, because a single daughter with brothers may be treated similarly to boys by parents, whereas having a brother for boys simply means more competition for limited resources. The "spillover effects" seem to apply for a girl with two brothers as well. Nevertheless, adding another girl begets "reference group effects," as observed in the negative sign, though of a smaller magnitude, for the group of a girl with one brother and one sister or with no brothers and one sister. A girl with no brothers and two sisters was worse off than boys with the same sibling composition due to absence of "spillover effects." We surmise that this was because parents without a male offspring in a multi-children household lacked the willingness to invest in their female children's health. While these results were as expected, the lack of statistical significance limits us from making stronger conclusions. This is a limitation in our study that arises from a lack of sample size for various sibling compositions especially in the one-child policy era in China. A larger sample size would have allowed us to examine sibling composition effects in greater detail.

Conclusion

In this paper, we investigated whether there was evidence of gender gap in health status in China during the less stringent years of the One-child policy. We examined whether only-child girls and boys fared better than children with siblings, particularly those with any brothers, and whether there was any difference in parental investments in child's health

[9] We controlled for birth order to tease out any birth order effects but found none.

between only-child daughters and only-child sons. Presumably, the girls examined were those who were wanted by their parents since those who were born survived sex-selective abortion, infanticide, and abandonment in childhood. Moreover, given the upward trend in economic growth and facilitated industrialization, gender gap in parent's investments should be diminishing. Therefore, if evidence of parents' gender bias found in health outcomes for daughters is credible, it would be a conservative estimate.

We found that being a female decreased height-for-age approximately by z-score 0.08. This meant that girls on average were 8.9 % shorter in height than boys. Further, having a male sibling decreased children's height-for-age by 0.17 z-score which amounted to a 19 % decrease in height.

On disaggregating our samples, there was no significant gender gap in health status between boys and girls in only-child households confirming our belief that girls born into only-child households were a "wanted" child. The fact that we did not see a significant impact of having a female sibling on the health of children but we saw a strong negative impact of a male sibling seems to suggest that the gender gap that we see happened not because of sibling rivalry for resources but because of a strong male preference amongst parents. We found this strong predilection for sons only in the rural areas. Further, this gender gap was absent in one-child only households in rural areas. However, in multiple-children households in rural areas, we found a wide gender gap of 10.6 % suggesting that rural girls with siblings were the worst off due to son preference. On further investigating the effect of sibling composition, we found that without female siblings, having a brother may have had a positive "spillover" effect on a girl's health status suggesting that girls as long as they are the only girl in the household were "wanted."

Clearly, China's one-child policy has had a tremendous influence on family structures, and that gender, number of siblings, and sibling composition played a central role in determining human capital accumulation for Chinese children. The coercive methods used by the enforcement of the policy with a neglect of the attitude of son preference has been problematic, since the resulting imbalanced sex ratio and "spoiled" children are projected to spawn politically and socially unstable factors and new problems in the Chinese society. Our results suggest that it is exigent for the government to form policy measures, such as pension for the elderly, health care programs and perhaps cash incentives to tackle parents' proclivity for sons over daughters. Neglecting girls' health, particularly in rural areas may well impose a tremendous long term cost to Chinese society and economy even as China abolishes its one-child policy. While we did not get to the birth order effects, it might be interesting to explore which boy in multiple boy households is worse off, especially now that parents will have multiple children after the one child policy has been relaxed. Another interesting question for the future is: Would income transfer to families in rural areas bring about a change in the deep rooted culture of son preference or would it be a structural change in the Chinese society that will force girls to be valued more?

References

Becker, G. S. (1960). An economic analysis of fertility. In *Demographic and Economic Change in Developed Countries*. Universities-National Bureau. Columbia University Press, pp. 209–240. Retrieved from http://www.nber.org/chapters/c2387.pdf.

Becker, G. S., & Lewis, H. G. (1973). On the interaction between the quantity and quality of children. *Journal of Political Economy, 81*(2), S279–288. Retrieved from http://www.jstor.org.opac.sfsu.edu/action/showPublication?journalCode=jpoliecon.

Bian, J. (1996). Parental monetary investments in children: A focus on China. *Journal of Family and Economic Issues, 17*(1), 113–139. doi:10.1007/BF02265033.

Bongaarts, J., & Greenhalgh, S. (1985). An alternative to the one-child policy in China. *Population and Development Review, 11*(4), 585–617. doi:10.2307/1973456.

Bredenkamp, C. (2009). Policy-related determinants of child nutritional status in China: The effect of only-child status and access to health care. *Social Science & Medicine, 69*, 1531–1538. doi:10.1016/j.socscimed.2009.08.017.

Bulte, E., Heerink, N., & Zhang, X. (2011). China's one-child policy and 'the mystery of missing women': Ethnic minorities and male-biased sex ratios. *Oxford Bulletin of Economics and Statistics, 73*(1), 21–39. doi:10.1111/j.1468-0084.2010.00601.x.

Cameron, L., Erkal, N., Gangadharan, L., & Meng, X. (2013). Little emperors: Behavioral impacts of China's one-child policy. *Science, 339*(6122), 953–957. doi:10.1126/science.1230221.

Ebenstein, A. (2011). Estimating a dynamic model of sex selection in China. *Demography, 48*(2), 783–811. doi:10.1007/s13524-011-0030-7.

Ebenstein, A. Y., & Sharygin, E. J. (2009). The consequences of the "missing girls" of China. *The World Bank Economic Review, 23*(3), 399–425. Retrieved from http://0-wber.oxfordjournals.org.opac.sfsu.edu/content/23/3/399.short.

Gao, M., & Yao, Y. (2006). Gender gaps in access to health care in rural China. *Economic Development and Cultural Change, 55*(1), 87–107. doi:10.1086/505720.

Garg, A., & Morduch, J. (1998). Sibling rivalry and the gender gap: Evidence from child health outcomes in Ghana. *Journal of Population Economics, 11*(4),471–493. Retrieved from https://www.nyu.edu/projects/morduch/documents/articles/1998-12-Sibling_Rivalry_and_Gender_Gap.pdf.

Guilmoto, C. Z. (2012). *Sex imbalances at birth: Current trends, consequences and policy implications*. Bangkok: UNFPA Asia and the Pacific Regional Office. Retrieved from http://www.unfpa.org/publications/sex-imbalances-birth.

Hill, R. C., Griffiths, W. E., & Lim, G. C. (2010). *Principles of Econometrics* (4th edn.). Hoboken, NJ: John Wiley & Sons, Inc.

Lee, M. (2012). The one-child policy and gender equality in education in China: Evidence from household data. *Journal of Family and Economic Issues, 33*(1), 41–52. doi:10.1007/s10834-011-9277-9.

Li, H., Yi, J., & Zhang, J. (2011). Estimating the effect of the one-child policy on the sex ratio imbalance in China: Identification based on the difference-in-differences. *Demography, 48*(4), 1535–1557. doi:10.1007/s13524-011-0055-y.

Li, J., & Cooney, R. S. (1993). Son preference and the one child policy in China: 1979–1988. *Population Research and Policy Review,12*(3),277–296. Retrieved from http://www.jstor.org/stable/40217602.

Millimet, D. L., & Wang, L. (2011). Is the Quantity-quality trade-off a trade-off for all, none, or some? *Economic Development and Cultural Change, 60*(1), 155–195. doi:10.1086/661216.

Murphy, R., Tao, R., & Lu, X. (2011). Son preference in rural China: Patrilineal families and socioeconomic change. *Population and Development Review, 37*(4), 665–690. doi:10.1111/j.1728-4457.2011.00452.x.

Osmani, S., & Sen, A. (2003). The hidden penalties of gender inequality: fetal origins of ill-health. *Economics and Human Biology, 1*(1), 105–121. doi:10.1016/S1570-677X(02)00006-0.

Report of a WHO expert committee (1995). *Physical status: The use and interpretation of anthropometry (WHO Technical Report Series No. 854)*. Geneva: World Health Organization. Retrieved from http://www.who.int/childgrowth/publications/physical_status/en/

Sen, A. (1990). More than 100 million women are missing. *The New York Review of Books*.

Short, S. E., Fengying, Z., Siyuan, X., & Mingliang, Y. (2001). China's one-child policy and the care of children: An analysis of qualitative and quantitative data. *Social Forces, 79*(3), 913–943. doi:10.1353/sof.2001.0025.

Short, S. E., & Zhai, F. (1998). Looking locally at China's one-child policy. *Studies in Family Planning, 29*(4), 373–387. doi:10.2307/172250.

Song, S., & Burgard, S. A. (2008). Does son preference influence children's growth in height? A comparative study of Chinese and Filipino children. *Population Studies, 62*(3), 305–320. doi:10.1080/00324720802313553.

Strauss, J. A., & Thomas, D. (1995). Human resources: Empirical modeling of household and family decisions In J. Behrman & T. N. Srinivasan (Eds.), *Handbook of development economics* (Vol. 3A, pp. 1883–2023). Amsterdam: Elsevier.

Strauss, J. A., & Thomas, D. (2007). Health over the life course In T. P. Schultz & J. Strauss (Eds.), *Handbook of Development Economics* (Vol. 4, pp. 3375–3474). Amsterdam: Elsevier.

Vidmar, S., Carlin, J., Hesketh, K., & Kole, T. (2004). Standardizing anthropometric measures in children and adolescents with new functions for egen. *The Stata Journal, 4*(1), 50–55. Retrieved from http://econpapers.repec.org/article/tsjstataj/v_3a4_3ay_3a2004_3ai_3a1_3ap_3a50-55.htm.

Wooldridge, J. M. (2010). *Econometric analysis of cross section and panel data* (2nd edn.). Cambridge, MA: MIT Press.

Wu, X., & Li, L. (2012). Family size and maternal health: Evidence from the one-child policy in China. *Journal of Population Economics, 25*(4), 1341–1364. doi:10.1007/s00148-011-0361-0.

Yu, M., & Sarri, R. (1997). Women's health status and gender inequality in China. *Social Science & Medicine, 45*(12), 1885–1898. doi:10.1016/S0277-9536(97)00127-5.

Zhai, F., & Gao, Q. (2010). Center-based care in the context of one-child policy in China: Do child gender and siblings matter? *Population Research and Policy Review, 29*(5), 745–774. doi:10.1007/s11113-009-9171-4.

Zhang, J., Xu, P., & Liu, F. (2016). One-child policy and childhood obesity. *China Economic Review*. doi:10.1016/j.chieco.2016.05.003.

Zhu, W. X., Lu, L., & Hesketh, T. (2009). China's excess males, sex selective abortion, and one child policy: Analysis of data from 2005 national intercensus survey. *The BMJ, 338*(1211). doi:10.1136/bmj.b1211.

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