

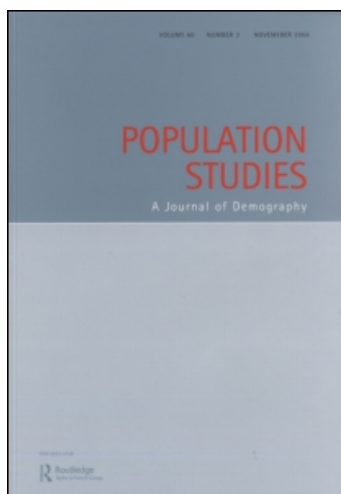
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Does son preference influence children's growth in height? A comparative study of Chinese and Filipino children

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Research has demonstrated that son preference has a serious impact on the survival and well-being of female infants and children in some parts of South and East Asia, but little is known about the consequences of son preference in later childhood and adolescence. We compare children's growth trajectories in height over childhood and adolescence in China, where the level of son preference is relatively high, and the Philippines, where it is relatively low. Children's height reflects long-term nutritional status and exposure to infectious diseases, both influenced by household decision-making and, presumably, by a preference for sons. Using data from two high-quality longitudinal studies and multilevel growth models, we find that male children in China show an additional height advantage relative to their female counterparts, when compared to the sex difference in growth trajectories in the Philippines. Further analysis reveals that the additional advantage of males in China is stronger in rural areas.

Keywords: son preference; gender inequality; child height; growth trajectory; China; Philippines

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Introduction

Research suggests that son preference has a serious impact on the health and well-being of female infants and children in some parts of South and East Asia, and is a cause of severe excess infant and child mortality among females (Sudha and Rajan 1999; Croll 2000; Bandyopadhyay 2003). These societies often show skewed population sex ratios strongly favouring males (Gu and Roy 1995; Scharping 2003; Cai and Lavelly 2004), in contrast to countries without strong traditions of son preference, where the biological advantage of females is evident after infancy. Relatively little is known, however, about the extent to which son preference influences the well-being of children who survive infancy (Pande 2003). Some have argued that parents with a preference for sons discriminate against daughters when they distribute scarce goods such as food or health care, leading to health, educational, economic, and reproductive difficulties later in life for these girls (Das Gupta 1987; Pebley and Amin 1991; Li 2004). However, several recent analyses have questioned whether

such discriminatory treatment towards surviving girls actually occurs, in the light of evidence that shows no preference for males in breastfeeding, immunization, and diarrhoeal disease treatment-seeking behaviours of parents (Obermeyer and Cardenas 1997) or in the distribution of calories within the household, measured by caloric intake or anthropometric measurements (DeRose et al. 2000; Marcoux 2002). While researchers have begun to question the validity of policy interventions intended to address the disadvantages of female children, given that evidence for discrimination has not been found in recent studies, there is still limited evidence for or against discrimination toward girls who survive infancy and early childhood. Recent studies assessing sex differences in children's physical growth have used data sources such as the Demographic and Health Surveys (Sommerfelt and Arnold 1998; Burgard 2002; Marcoux 2002), which provide wide global coverage of children under the age of 5 but very limited information about older children, and no possibility of longitudinal follow-up of individual children as they grow.

In the analysis presented below we add to the evidence about the potential consequences of sex discrimination for the well-being of surviving female children by isolating the impact of son preference on children's growth in height during adolescence. Achieved height is a sensitive measure of health in childhood that reflects the interaction of genetic potential for height with nutrition, exposure to infectious diseases, and access to medical facilities (Martorell and Habicht 1986; Eveleth and Tanner 1990; Pelletier 1998; Alter 2004). The latter resources are sensitive to intra-household resource allocation and, presumably, to a preference for sons. We compare China, a country noted for its long and strong tradition of son preference, with the Philippines, used here to represent a similarly middle-income country environment in the same global region, but one lacking a strong tradition of son preference. We compare the growth in height of boys and girls as they age, using high-quality longitudinal data and controlling for some of the key household characteristics that could influence both inter-household and intra-household allocation of family resources. This analysis is among the first to provide evidence for an advantage of males in physical growth among Chinese children after infancy and early childhood, and suggests that son preference may continue to exert negative effects on girl children even among those who survive strong forces of selection in early life.

Gender inequality and children's height

We use the relative magnitude of the sex difference in height to measure son preference instead of more direct measures of sex-selective abortion and infanticide because these other behaviours are illegal in most societies, making measurement extremely difficult. While it can be challenging to obtain honest answers about how male and female children are treated by parents in a survey (Merli and Raftery 2000), child anthropometric status, such as measured height, is less controversial. Parents need not report any stigmatized behaviours and anthropometric measurement errors can be minimized, especially when repeated height measurements are taken.

Son preference could generate inequality between the sexes in children's growth in height through discriminatory resource allocation within the household: parents may prioritize investing in sons, and be more willing to compromise the well-being of daughters whenever there is the possibility

of resource competition or dilution. For example, Graham et al. (1998) report that Chinese girls in Anhui province were breastfed for significantly shorter periods than boys, especially higher-parity girls, while Short et al. (2001) report that Chinese girls in general received less care from parents than boys, and even less in places where the one-child policy was strict. A study of schoolchildren in a schistosomiasis-endemic area of China showed that greater growth retardation among girls is likely to be due to traditional norms that mean that girls spend more time in schistosomiasis-infested environments and receive less when food is distributed (Zhou et al. 2005). By contrast, studies of Filipino children do not point towards differential treatment by sex. Studies have shown that males are more likely than females to be stunted, or height-for-age deficient, in the first year of life, while females are more likely to be stunted in the second year (Adair and Guilkey 1997). On the other hand, girls stunted at age 2 years are more likely to experience catch-up growth than their male counterparts (Adair 1999).

Despite conditions that would seem to disadvantage girls in countries with a strong tradition of son preference, recent large-scale comparative studies of children's growth across nations generally do not demonstrate an advantage for boys even in settings where preference for sons is strong. One study comparing cross-sectional Demographic and Health Survey data from 34 countries found little evidence of consistent disadvantage among females across three measures of nutritional status, though China was not among the countries included (Sommerfelt and Arnold 1998). However, in a similar study that did include data from China, the more recent Chinese samples did not show an advantage for males for children under age 6 (Marcoux 2002). Filipino boys were more likely than girls, or equally likely, to show growth retardation in height-for-age or weight-for-age or weight-for-height across five surveys covered in the same study (Marcoux 2002). These studies, however, are based on cross-sectional data that focus on relatively narrow age ranges and suffer from weaknesses that are inherent in all cross-sectional research, such as lack of adequate control for individual-level heterogeneity.

We take a different approach: instead of comparing the sex difference in children's heights at specific ages or historical moments, we use longitudinal anthropometric data and compare sex differences in growth trajectories over the entire period of childhood and adolescence (ages 0–17). Our strategy implicitly assumes that the level of son preference

can be measured at the country level, and that China represents a setting of strong son preference, while the Philippines represent a setting of little to no son preference. There are many other factors that differentiate the two countries, but we are most interested in the impact of son preference, a dimension on which these two cases differ substantially.

*Son preference in comparative perspective:
China and the Philippines*

Past research suggests that infanticide among females, sex-selective abortion, or (sometimes and) sex-selective birth misreporting are prevalent in some parts of China, especially in rural areas (Johansson and Nygren 1991; Zeng et al. 1993; Coale and Banister 1994; Ren 1995), and these behaviours are responsible for the increasingly unbalanced sex ratio at birth (the number of males born divided by the number of females born). The sex ratio at birth increased from 1.08 in the early 1980s to 1.12 in the early 1990s, then to 1.18 in the early 2000s (Gu and Roy 1995; Scharping 2003; Cai and Lavelly 2004). Similarly, the sex ratio of infant mortality (infant mortality for females divided by infant mortality for males) increased from 0.95 in the early 1980s to 1.15 in the early 1990s until the early 2000s (Attané 2005). The strong tradition of son preference in China dates back to the origins of ancestral worship thousands of years ago, when it was supported by the imperial state and Confucian ideology (Lee and Wang 1999). Despite the drastic social changes occurring during the past half century, preference for male children survives. The economic reform and decollectivization since the late 1970s have increased the value of rural children's labour, especially that of male children, and restored the central role of male household heads in decision-making and resource allocation (Li 2004), while also weakening the state's capacity to interfere in family issues (Davis and Harrell 1993). Furthermore, family planning policies intended to place strict limits on total family size and implemented since the late 1970s have, if anything, intensified sex discrimination at birth (Graham et al. 1998; Li et al. 2000; Chu 2001), though the strength and enforcement of these policies have changed over time and varied from place to place.

By contrast, the Philippines has long been characterized by a more egalitarian treatment of male and female children (Bautista 1988) and a greater sense of gender egalitarianism among adults, compared with

many East and South Asian countries including China (Hunt 1965; Upadhyay and Hindin 2005). More balanced sex ratios at birth in the Philippines reflect this difference (Mason 1987; Bautista 1988; Hudson and Boer 2004), as does the more typically higher child mortality among males than females (Martin et al. 1983; National Statistics Office [Philippines] and ORC Macro 2004). In light of these historical societal differences and the existing data on sex ratios at birth and differential infant mortality by sex, we expected to observe stronger evidence of son preference among children and adolescents in China than in the Philippines, as evidenced by their growth in height.

Hypothesis 1: The advantage in height of males is larger in China than in the Philippines.

Are rural and urban children affected equally?

It cannot be assumed that the expression and manifestation of son preference are the same for all social groups in China; the urban-rural division in life chances may be particularly important. The household registration or *hukou* system, an internal passport system designed to prevent rural-to-urban migration (Wu and Treiman 2004), ensures that urban Chinese residents have exclusive access to higher-income jobs, better nutritional and educational opportunities, and the state-sponsored pension and health care systems (Liu et al. 1999; Braveman and Tarimo 2002). While many urban parents can rely on the state for economic support and services as they age, in rural areas the responsibility for elderly parents falls to sons. The patrilineal family system in China thus creates a strong disincentive for rural Chinese parents to invest heavily in daughters, because they will not make significant financial contributions to their natal families as adults (Das Gupta et al. 2003). Moreover, in urban areas of China men and women are more equally rewarded at work, kin make up a smaller portion of social networks, and the differential impact of exogamy on old-age support of the parents of husbands and wives is minimized.

Furthermore, since basic necessities like food and medical care may be more easily accessible for urban residents, expressions of son preference may take a different form there than in rural areas, where parents may withhold these scarcer basic resources from girl children. Some research suggests that wealthier families (and by extension here, urban families) may choose sex-selective abortion to

achieve their family-formation preferences, rather than withholding resources from surviving female children (Das Gupta and Bhat 1997; Bhat and Zavier 2003; Dubuc and Coleman 2007). In particular, sex-selective abortion may be a more common response for urban than for rural Chinese families, because family-size limitation policies are more strictly enforced in urban areas (Short et al. 2001). These conditions suggest that the manifestation of son preference in sex differences in children's achieved height may be more evident in rural areas than in urban areas of China.

In contrast to the conditions prevailing in China, there is no *hukou* system in the Philippines restricting rural-to-urban migration, though life chances and living conditions certainly differ between urban and rural areas. Furthermore, a bilineal family system operates in the Philippines, with both sons and daughters expected to help parents in their old age. Investment in children reflects the more gender egalitarian pattern; Filipino parents tend to give land to sons because of the domination of men's labour in rice farming, but tend to invest more in the schooling of their daughters because returns to the schooling of females have risen in the non-agricultural sector (Quisumbing 1994; DeGraff et al. 1996). Given the different incentives associated with patrilineal and bilineal family systems, resource constraints in rural areas, and socio-historical conditions in China and the Philippines, we expected less variation in the sex difference in height between rural and urban areas in the Philippines than in China.

Hypothesis 2: The rural/urban difference in the advantage in height of males is greater in China than the Philippines.

Data and methods

Data

In order to compare patterns of children's growth in height between China and the Philippines, we use data from two longitudinal surveys: the China Health and Nutrition Survey (CHNS) and the Cebu Longitudinal Health and Nutrition Survey (CLHNS). The CHNS was conducted by the Carolina Population Center at the University of North Carolina at Chapel Hill, the National Institute of Nutrition and Food Safety, and the Chinese Center for Disease Control and Prevention. The CHNS is a panel survey with a multistage clustered sample of 3,800 households in nine Chinese provinces, yielding

a total of 16,000 selected individuals. Five waves of CHNS data are publicly available, collected in 1989, 1991, 1993, 1997, and 2000, and measures of height for children are available at all waves.

The CLHNS was conducted by the Carolina Population Center at the University of North Carolina at Chapel Hill, the Nutrition Center of the Philippines, and the Office of Population Studies at the University of San Carlos. A baseline survey (1983–86) was conducted among 3,327 women living in 33 randomly selected communities from the Metropolitan Cebu area. Women in their sixth to seventh month of pregnancy were interviewed so that all impending births could be identified. The baseline sample included 3,080 singleton live births, and subsequent surveys took place immediately after birth and then at bimonthly intervals for 24 months. Three follow-up surveys were conducted in 1991–92 (mean age of children 8 years, 74 per cent of original sample), 1994–95 (mean age 11.5 years, 71 per cent of original sample), and 1998 (mean age 15.5 years, 68 per cent of original sample).

For the purposes of our analysis we pooled the two data-sets and generated a dichotomous variable so that 1 = Chinese respondent and 0 = Filipino respondent. Figures 1 and 2 plot children's height by age in years for the CLHNS and CHNS samples, respectively. Note that these figures do not distinguish the survey wave from which measures were drawn—observations at a given age could be drawn from any survey wave for the Chinese sample, while the cohort structure of the Filipino sample means that most observations at a given age were drawn at the same survey wave. Clear differences in data structure are thus apparent when comparing these figures; data collection followed a cohort of births in the CLHNS with varying intervals between measurements over time, while the CHNS was a period-oriented survey of individuals of all ages in selected households.

Variables

To maintain maximum comparability between the two samples, we used a small number of essential variables in our analysis. Survey wave-specific summary statistics for selected variables from the CLHNS and the CHNS samples are displayed in Tables 1 and 2, respectively. The dependent variable was the selected child's height at a given age, operationalized first as raw height in centimetres, second as standardized height-for-age *z*-score based

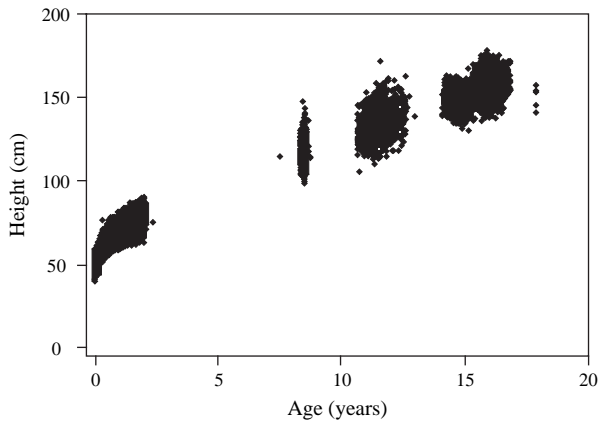


Figure 1 Distribution of Filipino children's height measurements by age

Source: As for Table 1

on the 1990 UK growth reference, and third as a dichotomous indicator of stunting, or a height-for-age z -score more than two standard deviations below the reference median. The child's age and sex were two important covariates in analyses using raw height measures, where age was treated as a time-varying variable. In both the CHNS and the CLHNS, a child's exact age was computed by subtracting his or her date of birth from the date of measurement in each data collection wave. Sex was measured at baseline with male = 1 and female = 0. Age and sex were implicitly controlled in z -score measures of height-for-age and stunting, which reflect an individual child's deviation from the reference population standard at a given age and sex.

Household characteristics explored include type of place of residence, with rural = 0 and urban = 1, mother's height in centimetres, and mother's education in years completed. The urban-rural distinction played an important role in our theoretical

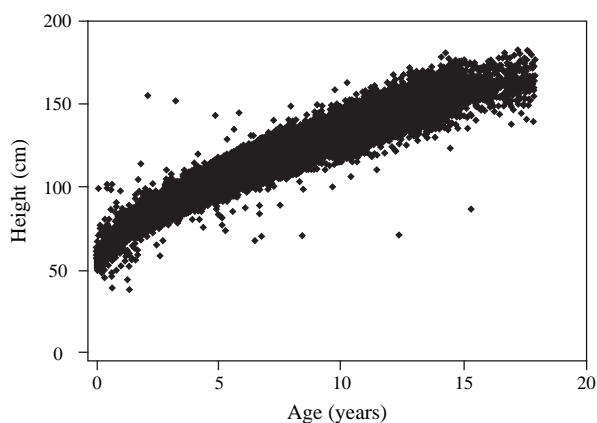


Figure 2 Distribution of Chinese children's height measurements by age

Source: As for Table 2

argument, and comparing Tables 1 and 2 reveals that owing to different sampling strategies the Filipino sample was predominantly urban (more than 70 per cent of children), while the Chinese sample was predominantly rural (about 70 per cent of children). Without adequate control for type of place of residence, the comparison between China and the Philippines was likely to be confounded by these different residential distributions. Mother's height was used to control for genetic differences in growth potential among children and also reflects her own socio-economic background, with taller mothers having experienced on average relatively better nutrition and health-producing resources in early life. Mother's education is an important indicator of children's socio-economic status, and more highly educated mothers may be able to provide better care in pursuit of their children's health and development (Thomas and Strauss 1992), but may also differentially support or practise actions that favour sons. We were unable to include further measures of household socio-economic status or wealth owing to data limitations inherent in the comparison of two very distinct panel studies; the potential consequences of this omission are considered in the discussion. Future studies should include such measures whenever possible, to better understand within-country and between-country differences in the expression and consequences of son preference and the way these may vary by household socio-economic status.

Outliers and missing data

Before embarking on the analysis, we addressed outlier observations present in the Chinese CHNS sample (shown in Figure 2). Ten extreme outlier observations were omitted to reduce bias. A further 81 cases were dropped from the Chinese sample in the process of computing the z -scores because their values were lower than -5 or greater than 5 . Such exclusions are typical when using z -scores for anthropometric measurements; these outliers are very likely to represent measurement or coding errors rather than abnormal growth patterns. We also experimented with other criteria to identify and exclude outliers; our results are quite robust to different exclusion rules.

Missing data in longitudinal studies arise for a number of reasons, ranging from panel attrition to respondents' refusal to answer certain questions, and including coding errors. The multilevel models

Table 1 Cross-wave comparison of selected variables in the Cebu Longitudinal Health and Nutrition Survey (CLHNS), 1983–98

	Average height (cm)	Average age (years)	% Male	% Urban	Average mother's education (years)	Average mother's height (cm)	Average standardized height-for-age z-scores	% Stunted
At birth	49.33	0.0	53	76	7.70	150.83	-0.68	13
2nd month	56.36	0.16	53	76	7.66	150.85	-0.41	6
4th month	61.11	0.33	53	75	7.64	150.83	-0.73	13
6th month	64.34	0.49	53	75	7.61	150.83	-1.00	16
8th month	66.85	0.66	53	75	7.60	150.83	-1.16	22
10th month	69.00	0.82	53	75	7.60	150.82	-1.32	28
12th month	70.80	0.99	53	75	7.60	150.86	-1.54	36
14th month	72.39	1.16	53	75	7.58	150.87	-1.73	42
16th month	73.78	1.32	53	75	7.58	150.82	-1.92	50
18th month	75.17	1.49	53	75	7.59	150.85	-2.07	52
20th month	76.55	1.66	53	75	7.58	150.84	-2.14	52
22nd month	77.79	1.82	53	75	7.57	150.86	-2.21	60
24th month	79.15	1.99	53	75	7.54	150.84	-2.23	56
1991	117.70	8.49	53	74	7.54	150.90	-2.23	61
1994	133.78	11.52	52	73	7.67	150.89	-1.82	45
1998	154.00	15.50	52	72	7.66	150.92	-2.00	49

Source: Cebu Longitudinal Health and Nutrition Survey, the Philippines, 1983–98.

used in our analysis perform well using longitudinal samples with missing data on the dependent variable (Hox 2002; Raudenbush and Bryk 2002). We used listwise deletion to deal with missing values in explanatory variables. We also conducted multiple imputation to assess the adequacy of the listwise deletion method. Results based on multiple imputation of data are highly consistent with results based on listwise deletion, and are available upon request.

Analytic approach

We used the multilevel model as our main analytic tool (Raudenbush and Bryk 2002; Goldstein 2003). Sex difference in individual growth trajectories can be compared across countries using the following model framework:

$$Y_{it} = \pi_{0i} + \pi_{1i}A_t + r_{it} \quad (1)$$

$$\pi_{0i} = \beta_{00} + \beta_{01}M_i + \beta_{02}C + \beta_{03}M \cdot C + \mu_{0i} \quad (2)$$

$$\pi_{1i} = \beta_{10} + \beta_{11}M_i + \beta_{12}C + \beta_{13}M \cdot C + \mu_{1i} \quad (3)$$

where Y equals child i 's height in centimetres at time t , A denotes age, M denotes male, and C denotes China. Equations (1)–(3) constitute a two-level linear growth model (Willett 1997; Singer and Willett 2003). In equation (1), the growth trajectory of child i throughout the period of observation is modelled as a linear function of age. Differences between individuals in growth trajectory are modelled in

equations (2) and (3) as a function of two time-constant covariates, male and China, reflecting the child's sex and nationality. To account for the fact that very few Chinese children have a height measurement taken at birth because of the study's sampling design, we centred the level-1 variable for age so that the intercept in equation (1) represents the i th child's height at age 2.

Several aspects of the above model are worth mentioning. First of all, unlike *post hoc* adjustment methods that treat the clustering of multiple observations per individual as a nuisance, multi-level models explicitly model the clustering effect in the form of random variance components. Multilevel models thus produce correct standard errors and also distinguish within-cluster variance from between-cluster variance, which allows us to understand how much variation occurs in an individual's growth pattern over time, as distinct from the variation that occurs between individuals. Second, although the true pattern of growth in height can be more accurately approximated using non-parametric or semi-parametric smoothing techniques, the parameters from those models are more difficult to interpret than those obtained from the linear growth model (Raudenbush and Bryk 2002). Furthermore, our specific interest in the comparison of heights of males and females at a given age makes it difficult to find a single ideal 'shape' for adolescents' growth patterns, because boys and girls go through typical growth spurts at different

Table 2 Cross-wave comparison of selected variables in the China Health and Nutrition Survey (CHNS), 1989–2000

	Average height (cm)	Average age (years)	% Male	% Urban	Average mother's education (years)	Average mother's height (cm)	Average standardized height-for-age z-scores	% Stunted
1989	89.79	2.71	53	30	6.85	155.43	-1.32	32
1991	100.83	4.12	53	29	6.75	155.32	-1.37	32
1993	109.90	5.53	53	29	6.72	155.36	-1.25	27
1997	125.77	8.16	53	29	6.98	155.69	-0.99	20
2000	134.94	9.75	53	29	7.37	156.04	-0.84	16

Source: China Health and Nutrition Survey, 1989–2000.

ages. Given these constraints, and the differences in data collection for the Chinese and Filipino samples that resulted in very different data structures (illustrated in Figures 1 and 2), we pursued the parsimonious approach of the linear growth model. Fortunately, the linear growth model performs well even when the growth in height shows mild departures from linearity, analogous to Ordinary Least Squares (OLS) regression in cross-sectional data analysis (Fox 1997).

Our first hypothesis was that the advantage in height of males is larger in China than in the Philippines. We tested this hypothesis by estimating two multilevel linear growth models: Model 1 is a simple linear growth model including only child's sex, nationality, and their interaction. Model 2 adds indicators of urban/rural place of residence, mother's education, and mother's height as further controls. To test our second hypothesis—that the rural/urban difference in the advantage in height of males is greater in China than the Philippines—we estimated Model 3, which includes all covariates in Model 2, plus two-way and three-way interaction terms between urban/rural place of residence, nationality, and child's sex.

We also estimated multilevel models using standardized z -scores and dichotomous indicators of stunting instead of raw height measurements, to provide results in a metric more common in anthropometric studies, but also to test the robustness of our results to a non-linear specification of age. We present these as subsidiary to the models using height in centimetres because the use of standardized z -scores in growth modelling has been criticized for eliminating growth information embedded in the raw measurements (Willett 1989, 1997). Specifically, raw height measures increase in magnitude over time, leading to a meaningful conceptualization of a growth trajectory, while z -scores can increase, decrease, or remain constant owing to their

implicit reference to a standard reference height for age. The z -scores have also been criticized for producing results that are sensitive to the choice of measurement metric (Seltzer et al. 1994; Frongillo 2004). In our case, we did not attempt to interpret the z -score models as growth models, but as random intercept models with multiple measurements per individual. In models using z -scores and indicators of stunting as outcomes, we divided the full age range into categories and investigated possible nonlinearities in the additional advantage of males in China as children age, which is not possible in a linear growth model. The z -scores model can be written as

$$Z_{ii} = \pi_{0i} + \pi_1 A_{2-9} + \pi_2 A_{10-18} + r_{ii} \quad (4)$$

$$\pi_{0i} = \beta_{00} + \beta_{01} M_i + \beta_{02} C + \beta_{03} M \cdot C + \mu_{0i} \quad (5)$$

$$\pi_{1i} = \beta_{10} + \beta_{11} M_i + \beta_{12} C + \beta_{13} M \cdot C \quad (6)$$

$$\pi_{2i} = \beta_{20} + \beta_{21} M_i + \beta_{22} C + \beta_{23} M \cdot C. \quad (7)$$

Model 4 in Table 5 is a two-level random intercept regression model using standardized z -score measurements, as depicted in equations (4)–(7). Replacing the OLS regression in equation (4) with a logistic regression, Model 5 estimates a two-level random intercept logistic regression using stunted status as the outcome measure.

Descriptive results

To inspect the unadjusted sex difference in height across countries, we calculated the mean difference in height for male and female children at every year of age in China and the Philippines, and then plotted the age-specific sex differences against age in Figure 3. Both lines in the figure have been smoothed using locally weighted regression (LOWESS smoothing) (Cleveland 1979). After adjusting the bandwidth of the lowess smoother to 2, the Chinese line has become very smooth but the

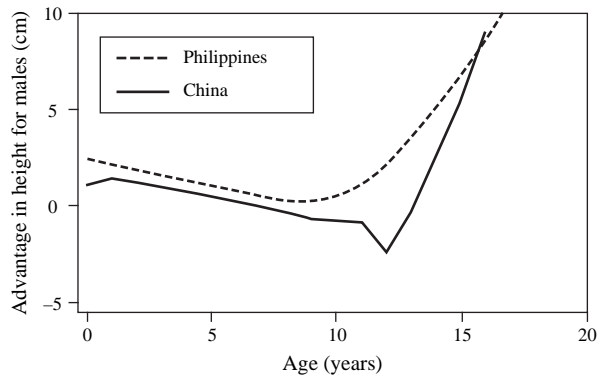


Figure 3 Lowest smoothed age-specific advantage for males in mean height for Chinese (CHNS) and Filipino (CLHNS) samples

Source: As for Tables 1 and 2

Filipino line remains more erratic, as a result of the data collection strategy discussed above. Figure 3 suggests that the overall advantage of males in children's height is larger in China than in the Philippines. Chinese boys are taller than Chinese girls at every age, while Filipino boys are shorter than Filipino girls between about age 8 and 13. The Filipino pattern approximates the typical sex difference in growth trajectories, in which an earlier growth spurt among girls leaves them taller than boys between the ages of about 10 and 13 years, with boys catching up and passing girls in average height thereafter (Abbassi 1998). Figure 3 provides important aggregate-level information that helps us understand the basic pattern of sex difference in children's growth in China and the Philippines. The pattern we see, however, is a mixture of between-individual variation and within-individual variation over time, so we turn next to multilevel models to partition these sources of variation.

Multivariate results

Is there an additional advantage for males in China?

Model 1 in Table 3 is the baseline linear growth model for children's height as they age; figures presented are coefficients with 95-per-cent confidence intervals in brackets. Coefficients for the 'initial status' component of the growth trajectory (equation (2)), shown in the top panel of Table 3, show that at age 2, Chinese girls are 9.4 centimetres taller than Filipino girls ('China' coefficient). As shown in Tables 1 and 2, Chinese children tend to be taller than their Filipino counterparts at every

age, and this height difference is reflected in Model 1 as the country-level difference among girls. Filipino boys are about 1.09 centimetres taller than Filipino girls ('Male' coefficient), while Chinese boys are about 0.86 centimetres taller than Chinese girls ($1.09 + (-0.23) = 0.86$), so that there is not a significant additional advantage for males in the Chinese sample at age 2 (China \times male: $\beta = -0.23$, 95 per cent CI = $-0.69, 0.23$). The 'rate of change' growth component of the model (equation (3)), seen in the middle panel of Table 3, shows that the linear growth rate for boys is 0.11 centimetres slower per year for boys than that for girls in the Philippines. By contrast, in China the linear growth rate for boys is 0.03 centimetres faster per year ($-0.11 + 0.15 = 0.03$) than that for girls at age 2. Finally, we observe an additional advantage for Chinese males of 0.15 centimetres in the rate of growth per year of age; over 16 years of growth this would sum to 2.4 centimetres, equal to about one-third of a standard deviation in the height distribution at age 16.

Adding controls for urban/rural place of residence, mother's education, and mother's height in Model 2 significantly changes the intercepts in both the initial status and the growth rate components from those of Model 1, but does not change the main findings of Table 3. The focal coefficients for Hypothesis 1, 'China \times male', continue to show no significant country-level difference in the advantage of males at age 2 ($\beta = -0.40$, 95-per-cent CI = $-0.86, 0.06$), while there remains an additional advantage for Chinese males in the growth rate ($\beta = 0.17$, 95 per cent CI = $0.09, 0.24$). The effects of these three added covariates are in keeping with expectation: urban children are taller ($\beta = 0.23$, 95-per-cent CI = $-0.00, 0.47$) than rural children at age 2, and grow significantly faster ($\beta = 0.10$, 95-per-cent CI = $0.07, 0.14$) than their rural counterparts. An additional year of mother's education is associated with a 0.15 centimetre increase in height at age 2, and a 0.02 centimetre increase in the linear growth rate per year. Children with taller mothers are significantly taller at age 2 and tend to grow significantly faster. Significant variance components for both initial status and growth rate in Models 1 and 2, shown in the bottom panel, indicate the presence of significant between-individual variation in the growth trajectory, controlling for the included covariates. The large individual-level residual variance component indicates that there is also significant within-individual variation in the growth trajectory. In sum, Models 1 and 2 show some support for our first hypothesis. Net of some key

Table 3 Coefficients from multilevel linear growth models of children's height in centimetres using Chinese CHNS and Filipino CLHNS samples

	Model 1: baseline model	Model 2: full model
<i>Initial status</i>		
Intercept	73.46*** [73.28, 73.64]	43.67*** [40.76, 46.58]
China	9.41*** [9.08, 9.75]	8.27*** [7.91, 8.63]
Male	1.09*** [0.84, 1.34]	1.17*** [0.92, 1.42]
China × Male	-0.23 [-0.69, 0.23]	-0.40 [-0.86, 0.06]
Urban residence		0.23 [-0.00, 0.47]
Mother's education		0.15*** [0.12, 0.18]
Mother's height		0.19*** [0.17, 0.21]
<i>Rate of change</i>		
Intercept	6.38*** [6.34, 6.41]	2.14*** [1.66, 2.62]
China	-0.31*** [-0.37, -0.26]	-0.32*** [-0.38, -0.26]
Male	-0.11*** [-0.16, -0.07]	-0.13*** [-0.17, -0.08]
China × Male	0.15*** [0.07, 0.22]	0.17*** [0.09, 0.24]
Urban		0.10*** [0.07, 0.14]
Mother's education		0.02*** [0.02, 0.03]
Mother's height		0.03*** [0.02, 0.03]
<i>Random effect variance</i>		
Child level: growth rate	0.22***	0.15***
Child level: initial status	9.35***	6.19***
Residual	28.38***	28.37***

* $p < 0.05$; ** $p < 0.01$; *** $p < 0.001$.

Note: 95% confidence intervals in brackets.

Source: As for Tables 1 and 2.

household characteristics and mother's height, an additional advantage for males in the growth rate persists in China throughout adolescence, compared with the sex difference in the Philippines.

Are rural and urban children affected equally?

Thus far, we have controlled for rural vs. urban place of residence because the Chinese and Filipino samples differ in the percentage drawn from urban areas. As argued above, the consequences of son preference for growth in height may also be stronger in rural areas of China than in urban areas

because of the combination of the distinct life chances faced by Chinese rural and urban families as a result of the *hukou* system, incentives embedded in the patrilineal family system, and the constraints rural families may face in obtaining food and medical care. In Table 4 we examine urban-rural differences in the additional height advantage of males, comparing children from China and the Philippines. A full model including all appropriate two-way and three-way interaction terms with urban residence is presented in the Appendix, but for ease of interpretation, in Table 4 we present linear combinations of the appropriate coefficients to show sex differences in the components of the

Table 4 Sex difference (advantage of males) estimated from multilevel linear growth model of children's height with urban-rural interactions, Chinese CHNS and Filipino CLHNS samples

	Coefficient	95% Confidence interval
<i>Initial status</i>		
Rural Philippines	1.01***	0.53, 1.49
Urban Philippines	1.23***	0.94, 1.52
Rural China	1.01***	0.56, 1.46
Urban China	-0.05	-0.99, 0.88
<i>Growth rate</i>		
Rural Philippines	-0.15***	-0.23, -0.07
Urban Philippines	-0.12***	-0.17, -0.07
Rural China	0.00	-0.07, 0.07
Urban China	0.11	-0.04, 0.26

* $p < 0.05$; ** $p < 0.01$; *** $p < 0.001$.

growth trajectory. Table 4 presents these sex differences and associated confidence intervals for four groups: rural and urban Filipinos and rural and urban Chinese respondents. At age 2 (initial status), rural Filipino boys are about 1.01 centimetres taller than rural Filipino girls, urban Filipino boys are 1.23 centimetres taller than their female counterparts, rural Chinese boys are 1.01 centimetres taller than rural Chinese girls, and urban Chinese boys are actually about 0.05 centimetres shorter than their female counterparts. For all but the final comparison, these sex differences are significantly different from zero. Of central interest here is that the rural/urban difference in the Philippines is about 0.22 centimetres ($1.01 - 1.23 = -0.22$), while it is a larger 1.06 centimetres in China ($1.01 - (-0.05) = 1.06$), a statistically significant difference (initial status coefficient for 'China \times male \times urban' in the Appendix).

The lower panel of Table 4 presents the same set of comparisons for the growth rate. Though they are taller at age 2, rural Filipino boys grow at a rate 0.15 centimetres slower per year than their female counterparts, and urban Filipino boys grow at a rate 0.12 centimetres a year slower, with both differences statistically significant. By contrast, there is no significant sex difference in the growth rate of rural or urban Chinese respondents. This suggests that while the sex gap in height may fall as Filipino children grow older, the disadvantage for rural females in China remains stable. In sum, the results in Table 4 show that Chinese boys have the expected height advantage at age 2 in rural areas, though there is no significant sex difference in urban areas in China. In the Filipino sample at

age 2, boys are taller than girls in urban and rural areas. As children grow older, the advantage of males among rural and urban Filipinos wanes, while the sex difference remains unchanged for Chinese young people, with no catch-up for females in rural areas.

Analyses using z-scores and stunting indicators

Turning to a final view of sex-differentiated patterns of growth, Table 5 reports results from two random intercept models that use standardized z-scores and indicators of stunting instead of the raw height measures used in Models 1-3. Model 4 is a random intercept linear regression model using z-scores as the dependent variable while Model 5 is a random intercept logistic regression using stunted status ($z\text{-score} < -2$) as the dependent variable. Rather than a linear effect for age, as shown in Models 1-3, Models 4 and 5 include dichotomous indicators for age groups (0-1 years (omitted category), 2-9 years, and 10-18 years) to explore potential age differences in any additional advantage in height of Chinese males, compared with the sex difference in the Philippines.

Results from Model 4 show that at ages 0-1, Chinese females have significantly higher z-scores (closer to the 'normal' value of 0) than Filipino females ($\beta = 0.21$, 95-per-cent CI = 0.11, 0.30), there is no sex difference in z-scores among the Filipino children ($\beta = 0.00$, 95-per-cent CI = -0.07, 0.08), and Chinese boys have slightly lower z-scores than Chinese girls ($\beta = -0.08$, 95-per-cent CI = -0.21, 0.04), but the difference is not statistically significant. Focusing for now on the within-country sex differences, we examine what occurs at older ages. At ages 2-9, Filipino boys fall slightly behind Filipino girls in average z-scores (Male \times Age 2-9: $\beta = -0.02$, 95-per-cent CI = -0.07, 0.04), but the difference is not statistically significant, while Chinese boys have significantly higher z-scores than Chinese girls ($-0.02 + 0.13 = 0.11$). At ages 10-18, Filipino boys show significantly lower z-scores than Filipino girls ($\beta = -0.15$, 95-per-cent CI = -0.21, -0.09), while Chinese males continue to show a stable advantage relative to Chinese females ($0.26 + (-0.15) = 0.11$). Figure 4 demonstrates this pattern of diverging differences in China and the Philippines visually. As children age, a small advantage to males in the Philippines shifts to a significant advantage to females, while in China a small advantage to females shifts to a significant

Table 5 Coefficients from two-level random intercept models of children's height using *z*-score and stunting indicators, Chinese CHNS and Filipino CLHNS samples

	Model 4: <i>z</i> -score model ¹	Model 5: stunting model ²
<i>Fixed effects</i>		
Intercept	-11.53*** [-12.19, -10.87]	22.87*** [20.86, 24.88]
China	0.21*** [0.11, 0.30]	-0.16 [-0.46, 0.15]
Male	0.00 [-0.07, 0.08]	-0.04 [-0.24, 0.15]
China × Male	-0.08 [-0.21, 0.04]	0.22 [-0.18, 0.62]
Age 2-9 ²	-0.85*** [-0.90, -0.81]	1.97*** [1.82, 2.12]
Age 10-18	-0.49*** [-0.53, -0.45]	0.94*** [0.81, 1.08]
Male × Age 2-9	-0.02 [-0.07, 0.04]	0.06 [-0.14, 0.26]
Male × Age 10-18	-0.15*** [-0.21, -0.09]	0.40*** [0.21, 0.59]
China × Age 2-9	0.54*** [0.46, 0.63]	-1.55*** [-1.85, -1.25]
China × Age 10-18	0.60*** [0.51, 0.70]	-1.92*** [-2.28, -1.56]
China × Male × Age 2-9	0.13* [0.01, 0.24]	-0.26 [-0.67, 0.15]
China × Male × Age 10-18	0.26*** [0.13, 0.40]	-0.55* [-1.05, -0.06]
Urban residence	0.12*** [0.07, 0.17]	-0.24** [-0.39, -0.09]
Mother's education	0.06*** [0.05, 0.06]	-0.14*** [-0.16, -0.12]
Mother's height	0.06*** [0.06, 0.07]	-0.15*** [-0.17, -0.14]
<i>Random effect variance</i>		
Child-level	0.66***	4.02***
Residual	0.71***	

¹The *z*-score is computed based on the 1990 UK growth reference.

²Stunting is defined as $z < -2$.

³The reference category for age group is age 0-1.

* $p < 0.05$; ** $p < 0.01$; *** $p < 0.001$.

Note: 95% confidence intervals in brackets.

Source: As for Tables 1 and 2.

advantage to males. The pattern of results for Model 5 is similar to that of the results for Model 4. Because the dependent variable in Model 5 is stunting status (1 = stunted, 0 = otherwise) rather than *z*-score (where a value closer to 1 is a good outcome), the signs of coefficients are the reverse of those obtained from Model 4. The key difference between Models 5 and 4 is that an advantage for Chinese males emerges only among respondents aged 10-18 when considering stunting, while it is apparent in both the 2-9 and the 10-18 age groups when examining *z*-scores.

Discussion

This study was designed to explore potential consequences of sex-based discrimination for the well-being of surviving female children by isolating any impact of son preference on children's growth in height over early childhood and adolescence. We compared China, a society known for its strong son preference tradition, with the Philippines, also a middle-income country in the same global region but without a strong tradition of son preference. Using both multilevel linear growth models with

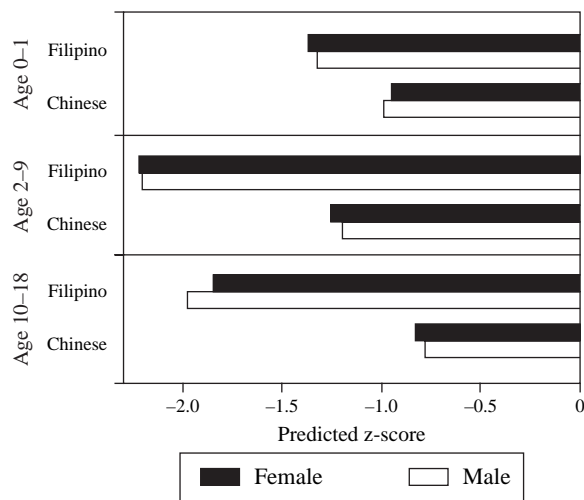


Figure 4 Predicted age-specific and sex-specific standardized z-scores for Chinese (CHNS) and Filipino (CLHNS) samples

Source: As for Tables 1 and 2

measures of actual height (Tables 3 and 4) and random intercept regression models with height-for-age z-score-based measures (Table 5) our results show considerable support for our hypotheses. First, we had argued that the advantage of males in height is larger in China than in the Philippines, and found this to be the case, though the Chinese advantage among males was evident in the growth rate and appeared only after early childhood. Our second hypothesis was that the rural/urban difference in the advantage in height of males is greater in China than the Philippines. Our results suggest that at age 2, rural females lag significantly behind rural males in China in height, while there is no sex difference in height in urban China, and that these gaps do not change much as children age. By contrast, there is no rural/urban difference in the Philippines in either the advantage of males in height at age 2, or the faster rate of growth among females as children age.

Our main result—evidence of an added advantage for males in children's growth in adolescence in China—conflicts with some findings in the recent literature. Other studies using a wide range of cross-sectional samples of children under the age of 5 have shown no advantage for males in children's height-for-age z-scores (Marcoux 2002) and no disadvantage in access to calories for female household members (DeRose et al. 2000). One potential explanation of the advantage found for males in this Chinese sample is our use of longitudinal data that cover a much wider age range—from birth to adolescence—whereas other studies are limited to

samples of young children. As our findings show, the advantage for Chinese males appears only after early childhood, and most strongly in the teenage years. We are also able to apply multilevel models to study the growth trajectories of individual children, rather than looking at relative height-for-age z-scores at a single point in time. The use of random effect models allows effective control for unmeasured individual heterogeneity and makes the separation of within-individual and between-individual variation possible. These are conventional reasons why results from longitudinal studies are preferred over those from cross-sectional studies. Longitudinal data and models allow us to isolate an advantage of males in settings with a strong son preference that has been difficult to find in previous studies.

More in keeping with existing theory and empirical findings, our results suggest that the negative effects of son preference on growth among girls may operate more strongly in rural areas of China. Studies in other contexts have shown stronger evidence of effects of son preference in rural areas; Clark (2000) reports that in India, rural women desire a higher proportion of sons than do urban women, and Ren (1995) notes relatively higher child mortality among female children in rural areas in China. Another study found that chronic energy malnutrition was present among rural Chinese women only, not urban women (Marcoux 2002). The present study may be the first to examine rural/urban differences in children's growth trajectories as they age, and suggests the need for further examination of differences in incentives for gender discrimination between rural and urban areas. Nonetheless, it is possible that the reason we do not observe significant sex differences in children's achieved height in urban areas in China is because urban families have easier access to the food and basic medical care necessary for normal growth—they are not scarce commodities to be shared unevenly between sons and daughters. Previous studies have shown that manifestations of son preference may vary by the socio-economic status of the family, with poorer families discriminating against girl children by withholding food or medical care while wealthier (or urban or wealthier and urban) families practise sex-selective abortion but meet the basic nutritional needs of surviving children (Das Gupta and Bhat 1997; Bhat and Zavier 2003; Dubuc and Coleman 2007). Without good information about the extent to which respondents in these samples had access to technology for sex-selection of pregnancies, it is difficult to estimate the nature or extent of the bias

that may affect our results owing to our limited controls for socio-economic status. Future research should look carefully at how the socio-economic status of the family and the availability of technology for sex-selection may condition the expression of son preference and its consequences for children.

The choice to directly compare the growth of children between societies with very different levels of son preference, from very strong (China) to low or nonexistent (Philippines) enhances our chances of observing the impact of son preference on child well-being. Nonetheless, there are at least two reasons to expect that our findings may still be a conservative estimate of the impact on their growth of discrimination against girl children. First, because of the availability of sex-selective abortion (and perhaps infanticide) in China (Ren 1995; Banister 2004), our Chinese sample may exclude the girls who would have been most disadvantaged had they survived gestation and early life. Those parents least interested in having female children and thus most likely to discriminate against them may have opted out of having daughters altogether (Goodkind 1996). Second, survey attrition as a result of child mortality in our samples, to the extent that it occurred, would probably have removed from the study those children in poorest health and potentially the female children experiencing the greatest levels of discrimination. Thus, our estimates of sex difference in children's growth in height in China may be somewhat conservative owing to differential sample selection by sex.

Some limitations of the present study should be considered when interpreting these results. Perhaps most importantly, we lack data on parents' behaviour toward male and female children, the mechanism purported to underlie our findings. We instead compare children's growth patterns to provide objective 'evidence' of the discriminatory treatment of girls through differential intra-household allocation of health-producing goods. While common to many studies of the consequences of sex preferences, this limitation inhibits our ability to make strong claims about the cause of the China-Philippines difference in the growth advantage of males among children and adolescents. In addition, different survey designs mean that the Chinese sample is dominated by those living in rural areas, where son preference is expected to have the strongest effects, while the Filipino sample includes a relatively small number of children living outside of an urban area. This means that our chances of observing son-preference effects in the Philippines may be relatively weaker than in China, where rural data are

more amply available. Despite these limitations, we feel that other strengths of the analysis make it a meaningful addition to the existing debate on sex preferences and their consequences.

Recent studies have suggested that despite common understandings of the impact of son preference on women's lives, and evidence in the form of skewed sex ratios at birth or local examples of differential care, convincing evidence for parental discrimination against girls is limited (Obermeyer and Cardenas 1997; Marcoux 2002). The present study, however, suggests that cultural preferences for sons do have consequences for the health and well-being of females in adolescence. Other examinations focusing on China (Das Gupta 2005), South Korea (Lee and Park 2006), and India (Clark 2000) have also reached the conclusion that cultural beliefs and preferences can influence parents' behaviour and have consequences for demographic outcomes. Taken together, such findings suggest that direct intervention to improve the conditions of girl children or to alter the incentives created by the patrilineal family system, restrictions on rural-to-urban migration, and the family-size limitation policies that prevail in China could have important consequences for women over the entire life course. We have shown that during the years crucial for educational attainment, itself a strong predictor of possibilities for later life achievement, girls (and particularly rural girls) suffer a growth handicap in China that is not present in the Philippines.

This study thus helps to bridge a gap in the current literature of demography and social stratification. One set of studies has shown that son preference influences infant survival among females and early life health and treatment (Graham et al. 1998; Short et al. 2001), while another set of studies has linked height deficits in early childhood to serious negative implications for school performance and labour market position in later life (Jamison 1986; Glewwe et al. 2001; Daniels and Adair 2004). The present study bridges the life-course gap between early childhood and later adolescence, to show that conditions of strong son preference may generate sex-specific growth trajectories that are particularly disadvantageous for rural girls, and which could have important implications across the life course. These findings help to create a more complete life-course picture of how gender inequality is produced and reproduced from life stage to life stage and across generations.

Notes

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- 2 Data from the Cebu Longitudinal Health and Nutrition Survey (<http://www.cpc.unc.edu/projects/cebu>) and the China Health and Nutrition Survey (<http://www.cpc.unc.edu/projects/china>) are publicly available from the University of North Carolina, Carolina Population Center (<http://www.cpc.unc.edu/>). An earlier version of this paper was presented at the 2006 annual meeting of the Population Association of America. The authors thank Donald J. Treiman, Barbara A. Anderson, and anonymous reviewers for useful comments.

Appendix A

Coefficients from multilevel linear growth model of children's height with urban-rural interactions, Chinese CHNS and Filipino CLHNS samples

	Model 3	
	Coefficient	95% CI
<i>Initial status</i>		
Intercept	43.92***	41.00, 46.84
China	7.73***	7.24, 8.22
Male	1.01***	0.53, 1.49
China × Male	0.00	-0.66, 0.66
Urban residence	-0.09	-0.50, 0.31
Male × Urban residence	0.23	-0.33, 0.78
China × Urban residence	1.42***	0.67, 2.17
China × Male × Urban residence	-1.06*	-2.10, -0.03
Mother's education	0.15***	0.12, 0.18
Mother's height	0.19***	0.17, 0.21
<i>Rate of change</i>		
Intercept	2.13***	1.65, 2.61
China	-0.27***	-0.35, -0.19
Male	-0.15***	-0.23, -0.07
China × Male	0.15**	0.04, 0.25
Urban residence	0.11**	0.04, 0.18
Male × Urban residence	0.03	-0.06, 0.12
China × Urban residence	-0.17**	-0.29, -0.05
China × Male × Urban residence	0.11	-0.05, 0.28
Mother's education	0.02***	0.02, 0.03
Mother's height	0.03***	0.02, 0.03
<i>Random effect variance</i>		
Child level: growth rate	0.14***	
Child level: initial status	6.19***	
Residual	28.36***	

* $p < 0.05$; ** $p < 0.01$; *** $p < 0.001$.

Source: As for Tables 1 and 2.

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