California Center for Population Research
University of California - Los Angeles

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Shige Song
Sarah A. Burgard

CCPR-018-07

July 2007

California Center for Population Research
On-Line Working Paper Series

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4 July 2007

Total Word Count: 11,342 (including abstract, text footnotes and references)
Number of Tables: 5 (plus 2 Appendix tables)
Number of Figures: 6
Running Head: Son Preference and Children's Growth

* Institute of Sociology, Chinese Academy of Social Sciences, E-mail: sgsong@ccpr.ucla.edu, Telephone: 86-10-64460722, Fax: 86-10-85195566.

** Departments of Sociology and Epidemiology and Population Studies Center, University of Michigan, E-mail: burgards@umich.edu, Telephone: 734-615-9538, Fax: 734-763-6887.

Acknowlegements: 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/). The authors would like to thank Don Treiman and Barbara Anderson for useful comments. Any remaining errors are our own.

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Abstract

A substantial body of 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 male and female 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 the level of son preference is relatively low. Children's height indexes well-being because it reflects long-term nutritional status and exposure to infectious diseases, both influenced by household decision-making and, presumably, by a preference for sons. Using longitudinal data from the China Health and Nutrition Survey and the Cebu Longitudinal Health and Nutrition Survey and multilevel growth models, we find that male children in China show an excess height advantage relative to their female counterparts, when compared to the male-female difference in growth trajectories in the Philippines. Further analysis reveals that the excess male advantage in China is evident only in rural areas. These findings are discussed in the context of the reproduction of gender inequality across the life course in high son preference settings.

Keywords: son preference, child height, growth trajectories, China, Philippines.

INTRODUCTION

Empirical 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. Sex differences in early life mortality exist in most of the developing world, with parts of South and East Asia showing particularly severe excess female infant and child mortality (Bandyopadhyay 2003; Croll 2000; Sudha and Rajan 1999). Population sex ratios strongly favoring males in these same regions also appear to provide evidence of son preference, as higher male child mortality in other societies eliminates the survival advantage males have at infancy, and results in sex ratios close to unity by young adulthood and even lower thereafter. Relatively little is known, however, about the extent to which son preference influences the well-being of children who survive infancy (Pande 2003). A generally held argument suggests that parents with a preference for sons discriminate against daughters when they distribute food and health care, leading to health, educational, economic, and reproductive difficulties later in life for these girls (Das Gupta 1987; Li 2004; Pebley and Amin 1991). However, several recent analyses have questioned whether such discriminatory treatment toward surviving girls actually occurs, based on evidence that shows no preference for males in breastfeeding, immunization, and diarrheal disease treatment-seeking behaviors of parents (Obermeyer and Cardenas 1997) or the distribution of calories within the household, measured by caloric intake or anthropometric measurements (DeRose, Das and Millman 2000; Marcoux 2002). Researchers have begun to question the validity of focusing policy attention or interventions to address disadvantages of female children that have not been found in recent studies, but there is still limited empirical support for either side of the debate when considering girls who survive infancy. Notably, recent studies examining children’s physical growth have used data sources such as the Demographic and Health Surveys (Burgard

2002; Marcoux 2002; Sommerfelt and Arnold 1998), which provide wide global coverage of children under the age of five but very limited information about older children, and no possibility of longitudinal follow-up of individual children as they grow.

In the present analysis we add to the evidence regarding potential consequences of gender discrimination for the well-being of surviving female children by isolating an impact of son preference on children's growth in height over early childhood and 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 (Alter 2004; Eveleth and Tanner 1990; Martorell and Habicht 1986; Pelletier 1998), all of which 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 height trajectories of boys and girls as they age through adolescence, using high quality longitudinal measurements and multilevel growth models, and controlling for key household characteristics that could influence both inter- and intra-household allocation of family resources. These longitudinal models allow for the decomposition of between- and within-individual variation in physical growth over time, providing a stronger test for an effect of gender discrimination among children as they age. To foreshadow our results, the present analysis is among the first to provide evidence for a male advantage 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.

Son Preference in Comparative Perspective: China and the Philippines

Past research suggests that female infanticide, gender selective abortion, and/or gender-selective birth misreporting are prevalent in some parts of China, especially in rural areas (Coale and Banister 1994; Johansson and Nygren 1991; Ren 1995; Zeng et al. 1993), and these behaviors are responsible for the increasingly unbalanced sex ratio at birth. China has a strong tradition of son preference (Lee and Wang 1999) dating back to the origins of ancestral worship thousands of years ago, when it was supported by the imperial state and Confucian ideology. 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 labor, especially male children, and restored the central role of male household heads in allocating resources (Li 2004), while also weakening the state's capacity to interfere in family decision-making (Davis and Harrell 1993). Furthermore, family planning policies to strictly limit total family size implemented since the late 1970s have, if anything, intensified gender discrimination at birth (Chu 2001; Graham, Larsen and Xu 1998; Li, Feldman and Li 2000).

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). More balanced sex ratios at birth in the Philippines reflect this difference (Bautista 1988; Mason 1987). As early as 1972, one study of married Filipino women's preferences for family size and composition showed that particularly in urban areas, a balanced sex composition was strongly desired, and that among those who favored an unbalanced pattern there was no noticeable preference for boys rather than girls (Stinner and Mader 1975).

Many investigators have identified traditional family structures as a root cause of societal differences in son preference (Das Gupta et al. 2003). The traditional bilineal family structure in the Philippines means that people expect both sons and daughters to help parents in their old age, such that birth order may have more consequence than sex roles, with older children bearing more of the responsibility (Hunt 1965; Stinner and Mader 1975). By contrast, the traditional Chinese family is patrilineal, and responsibility for elderly parents falls to the sons, leaving only a secondary role for birth order (Davis-Friedmann 1985; De Vos 1985; White 1987; Wolf 1984). In China, interdependencies between persons are generally confined to the stem family, and residually to the brothers of the same father, while in the Philippines such interdependencies can extend beyond the nuclear or stem family to neighbors or other kin, making elderly parents less dependent on children of either sex (De Vos 1985). Based on these differences, we expect that a child's sex will have greater consequences for his or her well-being in China than in the Philippines.

Differences across urban and rural areas are another important dimension of stratification systems in most developing countries, and interact with distinctive family systems to create different manifestations of son preference. Rural life is associated with labor-intensive agricultural work, strong kinship ties, and exogamy. In patrilineal family systems like those in China, these conditions create a strong disincentive for rural 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). In bilineal family systems like those in the Philippines, however, the situation is quite different. While Filipino parents tend to give land to sons because of the domination of male labor in rice farming, they tend to invest more in the schooling of their daughters because returns to female schooling have risen in the non-agricultural sector (DeGraff, Bilsborrow and Herrin 1996; Quisumbing 1994). Despite the abovementioned differences across family systems, however, urban residents in patrilineal and bilineal systems have more in common than do rural residents across these two family systems because urban life creates a new environment in which 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.

Importantly, the urban-rural division in life chances is more acute in China than elsewhere because of the household registration system, an internal passport system designed to prevent rural-to-urban migration (Wu and Treiman 2004). Urban residents have exclusive access to higher income jobs, better nutritional and educational opportunities, and the state-sponsored pension and health care systems (Braveman and Tarimo 2002; Liu, Hsiao and Eggleston 1999; Popkin, Lu and Zhai 2002), which reduce or eliminate the socioeconomic rationale for urban parents to prefer sons over daughters. Rural residents, on the other hand, live under conditions similar to those faced by their ancestors: labor-intensive farm work and a lack of reliable old-age support and health care, other than that which can be obtained from family or kin. Thus in rural areas of China, sons are perceived to be much more valuable than daughters and vital for the survival of the family. Given these conditions, we expect that the consequences of son preference for childrens' well-being will be noticeably stronger in rural areas than in urban areas.

Gender Inequality and Children's Height

Achieved height is a useful measure of children's health because it reflects their long-term nutritional status and exposure to infectious diseases, and also because it is simple to measure and less vulnerable to measurement errors than other indicators of well-being or treatment. In some studies it is difficult to ascertain the effects of son preference because actions like sex-selective abortion and infanticide are highly sensitive issues and illegal in most societies, including those with a skewed sex ratio. Therefore, it can be challenging to obtain honest answers about how sons and daughters are treated by parents in a survey setting (Merli and Raftery 2000). Child anthropometric status, by contrast, does not suffer from such problems; parents need not report any stigmatized behaviors and with proper training and careful supervision, measurement errors can be effectively controlled, especially when repeated height measurements are taken. Comparison of female and male children's growth patterns across countries with different systems of gender preference, after controlling for potential confounders, could be expected to reveal female children's long-term exposure to discrimination, if it exists.

How does son preference generate gender inequality in children's health and development? The key mechanism is gender-differentiated resource allocation within the household: parents prioritize investing in sons, and are more willing to compromise the well-being of daughters whenever there is the possibility of resource competition or dilution. In the most extreme cases, parents resort to sex-selective abortion or sex-selective infanticide to eliminate "unwanted" girls (Banister 2004; Hudson and Boer 2004; Ren 1995). Even among children who survive infancy, however, gender discrimination could influence health and may be reflected in growth patterns. For example, Graham and colleagues (1998) report that Chinese girls in Anhui province were breastfed for significantly shorter periods than boys, especially among higher-parity girls, while Short and colleagues (2001) report that Chinese girls in general received less care from parents than boys; they received even less care 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 girls spend more time in schistosomiasis-infested environments and receive less when food is distributed (Zhou, He and Ohtsuka 2005). By contrast, studies of Filipino children 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), but girls stunted at age two years are more likely to experience catch-up growth than their male counterparts (Adair 1999). There is some evidence from data collected in the late 1970s (Magnani et al. 1993) and late 1980s (Food and Nutrition Research Institute 1990) that before the age of five, Filipino boys showed a weight-for-age advantage over girls, but these differences in weight do not appear to have had long-term negative effects on female well-being, as gender differences in infant and child mortality favor girls in the Philippines, as is the case in other countries without notable patterns of son preference.

Despite conditions that would seem to disadvantage girls in countries with a strong son preference tradition, recent large-scale comparative studies of children's growth across nations generally do not demonstrate an advantage for boys. One study comparing cross-sectional Demographic and Health Survey data from 34 countries found little evidence for consistent female disadvantage 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 a male advantage for children under six (Marcoux 2002). Filipino boys were more likely than girls, or equally likely, to show growth retardation in height- 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 of control for individual-level hetetrogeneity. We take a different approach. Instead of comparing the gender difference in children's heights at specific ages or historical moments, we utilize high quality longitudinal anthropometric data and compare gender differences in growth trajectories over the entire period of childhood and adolescence (ages 0 to 17) in China and the Philippines. Our strategy implicitly assumes that the level of son preference can be ascertained at the country level, and that China represents a setting of high son preference, while the Philippines represent a setting of little to no son preference.[1] 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. Such a stark comparison may provide the necessary leverage to observe effects of gender discrimination on children's physical development.

Drawing upon existing theory and evidence, we aim to test the following research hypotheses:

Hypothesis 1: Based on existing literature suggesting a historically longer and stronger influence of son preference in China than in the Philippines, the male advantage in growth trajectories in height will be greater in China than in the Philippines. We refer to this as the "difference-in-difference hypothesis."

Hypothesis 2: Existing studies have suggested that gender discrimination should operate more strongly in rural areas than in urban areas. We hypothesize that the "difference-in-difference" in growth trajectories is greater in Chinese rural areas than in Chinese urban areas. Put differently, we expect the disparity between China and the Philippines in the gender difference in children's growth trajectories to be greater in rural areas than in urban areas.

[1] Using nationality to serve as a proxy measure for beliefs or behaviors that are difficult to measure in survey data is a common practice in international comparative research.

METHOD

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. The CHNS includes a household survey with measures of height for all residents, a community survey, a food market survey, and a health and family planning facility survey; we use the household survey in the present analysis.

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 (1983-1986) 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 living in 33 randomly-selected communities from the Metropolitan Cebu area were interviewed so that all impending births could be identified. The baseline sample included 3,080 non-twin live births, and subsequent surveys took place immediately after birth, then at bimonthly intervals for 24 months. Three follow-up surveys were conducted in 1991-1992 (mean age of children 8 years, 74% of original sample), 1994-1995 (mean age 11.5 years, 71% of original sample), and 1998 (mean age 15.5

years, 68% of original sample). The CLHNS collected individual, household, and community information; we use individual and household-level data.

For the purposes of this analysis we have pooled the two data sets together to create a master longitudinal data set, and generated a binary variable that equals one if a case belongs to the Chinese sample and zero if it belongs to the Filipino sample. Figure 1 and Figure 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 in years may 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 sample wave. Clear differences in data structure are 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 periodic survey of individuals of all ages in interview households.

Variables

To maintain maximum comparability between the two samples, we extract and utilize a small number of 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 main dependent variable is the physical measurement of the selected child's height; for the main multivariate analysis we use raw height measurements in centimeters, while in a sensitivity analysis that focuses exclusively on China we use standardized height-for-age *z*-scores based on the 1990 U.K. growth reference. The child's age and sex are two important covariates in analyses using raw height measures; age is time-varying – the measure changes from survey wave to wave. In both the CHNS and the CLHNS, a child's exact age is computed by subtracting his/her date of birth from the date of measurement in each data collection wave. Sex is a time-constant dichotomous variable measured at baseline with male = 1 and female = 0. Age and sex are implicitly controlled in z-score measures of height-for-age, which reflect an individual child's devation from the reference population height at a given age and sex.

Household characteristics explored here include type of place of residence, with rural = 0 and urban = 1, mother's height in centimeters, and mother's education in years. The urban-rural distinction plays an important role in our theoretical argument, and comparing Tables 1 and 2 reveals that while the Filipino sample is predominantly urban (more than 70% of children), the Chinese sample is predominantly rural (about 70% of children). Without adequate control for type of place of residence, the ostensible comparison between China and the Philipines is likely to be counfounded by a comparison between a largely rural and a largely urban sample of these populations. Mother's height is used to control for genetic differences in growth potential among children that may influence their physical growth, and reflects her own socioeconomic background, with taller mothers on average having experienced relatively better nutrition and health-producing resources in early life. Mother's education is an important indicator of children's socioeconomic 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).

To assess the robustness of our main analytic findings for Chinese children, the population for whom we expect to see the consequences of strong son preference, we also explore the influence of sibship composition, province of residence, and ethnic group of membership on our main results. These tests are conducted only for the Chinese sample because sibship information is unavailable in the Filipino study, and that study was conducted solely in the Cebu metropolitan area, while the Chinese sample was drawn from nine diverse provinces.

We explore the importance of ethnic group membership in the Chinese sample because family planning restrictions have been more lenient for ethnic minority group members than for those in the Han majority, potentially creating differential levels of incentive for gender discrimination. To explore the importance of sibship composition, we include three distinct sets of controls: the total number of siblings in the household (range: 0 – 6), the focal child's birth order (0 = first child, 1 = second child, 2 = third child, 3 = fourth and above), and two dichotomous indicators of the presence of an older male sibling (= 1) and/or an older female sibling (= 1). Net of all else, a "resource dilution" perspective on the consequences of sibship size would suggest that the greater the number of siblings in the household, the fewer resources that may be allocated to each one (Blake 1986). An alternative "competition" argument focuses on the effect of sibship order, and suggests that older siblings are allocated more resources than are younger ones. Finally, a few studies have shown that the sex composition of older siblings is influential for the focal child's well-being; in particular, younger girls with one or more older sisters may be particularly disadvantaged (e.g., Pande 2003). We test all of these theoretical possibilities to assess their impact on our results, but remain agnostic about the relative importance of these processes in the Chinese context.[2] Dichotomous variables are created to indicate the Chinese province of residence, and include indicators for Liaoning (11% of the sample), Helongjiang (11%), Jiangsu (10%), Shandong (8%), Henan (12%), Hubei (12%), Hunan (9%), Guangxi (13%), and Guizhou (14%). A dichotomous indicator is created to denote whether the focal child belongs to the Han majority ethnic group (= 0) or an ethnic minority group (= 1).

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[2] All sibship composition measures are based on the children who were present in the household at the time of the interview; siblings that are away at the time of the interview are not included in these measures.

Outliers and Missing Data

Figure 2 clearly shows that a fair number of outlier observations are present in the Chinese CNHS sample; to reduce the bias introduced by these outliers, we exclude ten observations with extreme values from the analysis.[3] A further 81 cases were dropped from the Chinese sample in the sensitivity analysis, for which models use height-for-age *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 most likely represent measurement or coding errors, rather than abnormal growth patterns. We have also experimented with other criteria to identify and exclude outliers; our results are quite robust to different exclusion rules.

Missing values in longitudinal studies arise for a number of reasons, ranging from panel attrition, to respondents' refusal to answer certain questions, to coding errors. In the present analysis we exclude observations that are missing height measurements, but include those with missing values on explanatory variables.[4] Multilevel models used in the present analysis perform well in longitudinal studies with missing data on the dependent variable (Hox 2002; Raudenbush and Bryk 2002). Various methods are used to deal with the difficulties that can arise when there are missing values on explanatory variables. Listwise deletion is the conventional method, and

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[3] Observations are excluded based on the following *ad hoc* rules: 1) child is reported to be 40 centimeters or shorter; 2) child is three years or older but is reported to measure 75 centimeters or less; 3) child is ten years or older but is reported to measure 100 centimeters or less; 4) child is younger than three but is reported to measure 140 centimeters or more; 5) child is younger than five but is reported to measure 150 centimeters or more.

[4] Approximately 1.5% of respondents are missing values on mother's height, and about 18.8% are missing values on mother's education. No other covariates have missing values.

yields unbiased estimates when data on explanatory variables is missing completely at random (MCAR).[5] Since MCAR is often too strong an assumption when using observational data, multiple imputation of missing values based on other covariates may yield better results. This is the second strategy we pursue here because the assumption that explanatory variables are missing at random (MAR) is often more reasonable than the MCAR assumption (Allison 2001; Rubin 1987) and recommends multiple imputation strategies. Nonetheless, in multivariate regression models if missing values on explanatory variables fit neither MCAR nor MAR assumptions, the listwise deletion approach may be preferable to more sophisticated approaches like multiple imputation (Allison 2001). We compare the results obtained using the traditional listwise deletion method versus multiple imputation using the method of chained equations (Raghunathan et al. 2001), as implemented in Stata (Royston 2004). Since multiple imputation does not involve calculation of a likelihood function for the data, and applying Rubin's rule to the vector of parameter estimates and their associated variance-covariance matrices does not work reliably (Li, Raghunathan and Rubin 1991), likelihood-based model fit indices are not reported for models based on multiply imputed data sets. We use results from listwise deletion models for model selection and model comparison, while using results from multiple imputation models as a check against potential biases introduced by the listwise deletion operation itself.

[5] Missing completely at random (MCAR) refers to the situation when the probability of having a missing value on variable *Y* is unrelated to the values of any observed variables in the data set. Missing at random (MAR) refers to the situation where the probability of having a missing value on variable *Y* may be related to other observed variables (which should thus be included as controls in multivariate models), but is unrelated to unobserved variables.

When results are consistent across specifications, we have greater confidence in the findings; when they are inconsistent, this suggests more caution in interpretation.

Analytic Approach

We use the multilevel model as our main analytic tool (Goldstein 2003; Raudenbush and Bryk 2002), estimating two different sets of multilevel models for the main and sensitivity analyses. Our research hypotheses concern comparisons between China and the Philippines, and can be tested using the following model framework:

$$Y_{ti} = \pi_{0i} + \pi_{1i} AGE_t + \pi_{2i} AGE_t^2 + r_{ti} \quad (1)$$

$$\pi_{0i} = \beta_{00} + \beta_{01} MALE_i + \beta_{02} CHINA + \beta_{03} MALE \cdot CHINA + \mu_{0i} \quad (2)$$

$$\pi_{1i} = \beta_{10} + \beta_{11} MALE_i + \beta_{12} CHINA + \beta_{13} MALE \cdot CHINA + \mu_{1i} \quad (3)$$

$$\pi_{2i} = \beta_{20} + \beta_{21} MALE_i + \beta_{22} CHINA + \beta_{23} MALE \cdot CHINA \quad (4)$$

Equations (1) – (4) constitute a simple yet powerful analytical framework, often referred to as the individual growth model (Singer and Willett 2003; Willett 1997). In Equation (1), the growth trajectory in height of child i throughout the period of observation is modeled as quadratic function of *AGE* and *AGE*-squared, his or her age in years at each survey wave.[6] Functional forms other than the quadratic can also be used; higher order polynomials usually achieve a better approximation of the true growth trajectory but result in estimated model parameters that

[6] We do not explore the possibility of both age and period variations in children's growth profiles because there is only a single birth cohort represented in the Filipino sample, and the Chinese sample only spans about 11 years; this means that the oldest respondents will age from six to 17 years over the course of observation, while the youngest will age from birth to 11 years, not a substantial enough overlap to explore.

are difficult to interpret, due to the added model complexity. Explatory analyses indicate that the quadratic model provides a good balance between accuracy and parsimony in the present analyses. Differences between individuals in growth trajectories are modeled in Equations (2) to (4) as a function of two time-constant covariates, $MALE$ and $CHINA$, reflecting the child's sex and nationality.[7] To account for the fact that very few Chinese children have a height measurement taken at birth because of that study's sampling design, we center the level-1 variable AGE so that the intercept in Equation (1) represents the i^{th} child's height at age two years.

Our first research hypothesis states that the male advantage in growth in height over childhood and adolescence is greater in China than in the Philippines. We test this "difference-in-difference" hypothesis by estimating four variations of the multilevel growth model described above: Model 1 is a simple linear model that excludes the age-squared term from Equation (1) and thus drops Equation (4), to explore a linear specification for growth as children age. Model 2 is presented in equations (1) through (4) above, and incorporates the quadratic component for age. Model 3 adds explanatory variables to equations (2) through (4) for rural or urban place of

[7] We include random effects in equations (2) and (3) to model the intercept, or initial status in height across children, and the linear or instantaneous rate of change, but do not include a random effect in equation (4) for the curvature component of the model because of limitations of the available Chinese data. To estimate a random effect for equation (3), three observations per child, on average, are required. To do so for the quadratic component represented by equation (4), at least four observations per child are necessary. There are too few children with four or more available observations to obtain a stable estimate of the random effect for equation (4), so we omit it.

residence, mother's education, and mother's height, and Model 4 is a re-estimation of Model 3 using multiply imputed data instead of listwise deletion of cases missing data on explanatory variables. To test our second hypothesis, that the "difference-in-difference" is greater in rural areas than in urban areas, we also estimate four models: Model 5 is the quadratic model outlined in equations (1) through (4) for rural children only, while Model 6 re-estimates Model 5 using multiply imputed data instead of listwise deletion. Models 7 and 8 are used to estimate the quadratic model for urban children only using listwise deletion and multiply imputed data, respectively.

Conducting sensitivity analyses for the Chinese sample only, as discussed above, requires a different analytic strategy. Information about sibship composition is available for the Chinese study only, and province and ethnic group differences are only relevant in the Chinese case for our purposes, so we cannot use the same difference-in-difference approach to directly compare children's growth in China and the Philippines. In order to arrive at a comparison of male and female childrens' relative growth profiles without the presence of explict reference groups from another social context, we use age- and sex-standized *z* scores instead of raw height measurements.[8] The use of *z* scores means that we are implicitly comparing the growth patterns of Chinese children with those of their same-age-and-sex counterparts in the 1990 U.K. reference population, and thus can comment on whether the deviation of Chinese children's heights from those in the reference population is greater among male or female children. These models also

[8] It is worth noting that even though age- and sex-standardized *z* scores have been widely used in empirical research on children's growth, they have been critiqued for distorting growth information embedded in the raw measurements (Willett 1997), and for producing results that are sensitive to the choice of measurement metric (Seltzer, Frank and Bryk 1994).

provide a useful basis for comparison with previous research employing *z* scores rather than raw height measurements, and demonstrate the relative severity of height-for-age deficiencies among Chinese children more clearly than those using raw heights. The models for sensitivity analyses take the following form:

$$Z_{ti} = \pi_{0i} + \pi_{1i} AGE_t + r_{ti} \tag{5}$$

$$\pi_{0i} = \beta_{00} + \beta_{01} MALE_i + \beta_{02} COVARIATES + \beta_{03} MALE \cdot COVARIATES + \mu_{0i} \tag{6}$$

Where Z_{ti} represents the age- and sex-standardized height *z* score for child i at time t. Compared to the first set of models, we have replaced the raw height measurements from equations (1) and (2) with age- and sex-standardized height-for-age *z* scores here in equations (5) and (6) and added *COVARIATES* one at a time denoting, respectively: (a) the sibship composition of the index child's household, (b) province of residence, and (c) ethnic minority group membership (each discussed below). We do not interpret this second set of models as multilevel growth models (note the absence of corollaries for equations (3) and (4) above), but simply as random intercept models with multiple measurements per individual. In other words, each child has an initial height π_{0j} that differs from other children's initial heights as a product of measured factors (e.g., sex) and random factors (random term μ_{0i}), but we do not model the trajectory of growth with age as a function of these non-random and random factors. In these models, age is simply a control variable. Even though the process of *z* score calculation has taken the age and sex of the child into consideration, including an indicator for age in the model captures any systematic deviation of the Chinese population from the U.K. reference population as children age. There is no equation to model π_{1i} and no need for an age-squared term. We also include an indicator for sex in the model to explore the excess male advantage in physical growth in China, the focus of the present study.

We begin the sensitivity analysis by estimating Model 9, which represents a baseline random intercept model as depicted in equations (5) and (6), but without controls for any covariates. We compare these results with Model 10, which adds the first control for sibship composition: the number of siblings in the household, and Model 11, which also adds an interaction term between the number of siblings and the child's sex. A finding of significant additive effects of sibship composition in Model 10 would be substantively interesting, but would not alter our main conclusions if the male advantage was still apparent. If the interaction between number of siblings and the focal child's sex improved model fit significantly, however, it could help to explain the excess male advantage in growth for Chinese males. Models 12 through 19 reflect the same general strategy as for Models 10 and 11, with a first model examining the additive effects of the other explanatory variables (presence of older female or male sibling, sibship order, province of residence, or ethnicity of origin), and a second model assessing the significance of interactions of these factors with the focal child's sex. Model 20 adds additive indicators of both province and ethnicity to Model 9. We compare AIC and BIC values to determine whether adjusting for these additional factors significantly improves model fit.

ANALYSIS

Descriptive Analysis

Since our goal is to compare the gender difference in height across countries, we calculated the mean gender difference in height for each age group in both samples, and then plotted the two series of age-specific differences against age in Figure 3.[9] 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 Filipino line remains more erratic, due to the data collection strategy discussed above. Figure 3 suggests that the overall male advantage 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 eight and age 13. The Filipino pattern approximates the typical gender difference in growth trajectories, where an earlier growth spurt among girls leaves them taller than boys between the ages of about 10 to 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 gender difference in children's

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[9] As adjunct information for Figure 3, Appendixes A and B report gender- and age-specific summary statistics for children's height using the Filipino CLHNS and Chinese CHNS samples. These summary statistics are based on the child's age in years rather than survey wave of data collection, making it possible that multiple data waves were used in calculating summary statistics for an age group. Boys and girls were interviewed separately in the final wave of the CLHNS; girls were interviewed in 1998 and 1999, while boys were interviewed during 1999 and 2000. This is reflected in Appendix Table 1: there are 64 girls at age 16 and zero at age 17, while there are only 43 boys at age 15 and zero at age 14. This peculiar data structure does not pose special difficulties for the multivariate analysis because multilevel growth models handle complicated data structures well (Raudenbush 2001). The distribution of observations among different age groups is more consistent in the Chinese sample, where observations from multiple data waves are frequently used to calculate summary statistics for a given age group.

growth in China and the Philippines. The pattern we see, however, is a mixture of between-individual variation and within-individual variation over time. Next we disentangle these two sources of variation using multilevel models.

Results from Multilevel Models

Testing the "Difference-in-Difference" Hypothesis

Model 1 in Table 3 is the baseline linear growth model for children's growth in height as they age. The linear growth model is appealing to many researchers because of its intuitive intepretation; specifically, the growth trajectory can be decomposed into an "initial status" component and a "linear growth rate" component. The "initial status" component of the model shows that at age two Filipino boys are about 1.09 centimeters taller than Filipino girls and Chinese boys are about 1.09 - 0.23 = 0.86 centimeters taller than Chinese girls; but Chinese boys do not show a statisically significantly greater advantage than Filipino boys at conventional levels of significance when each group is compared to their female counterparts (China × Male interaction, $\beta = -0.23$, $p > 0.05$). The "rate of change" or growth component of the model shows that the linear growth rate for boys is 0.11 centimeter per year lower than that for girls in the Philippines; but in China, the linear growth rate for boys is 0.04 centimeters per year greater (= -0.11 + 0.15) than that for girls at age two; the difference-in-difference in the linear growth rate at age two is statistically significant (China × Male interaction, $\beta=0.15$, $p < 0.05$).[10]

[10] We estimated three variance components for each model, as shown at the bottom of Table 3: the initial status variance, the linear rate of change variance, and the residual variance. The significant initial status and linear rate of change variances in all models show that individual

Adding quadratic components to the model improves the model fit, as indicated by favorable changes in both AIC and BIC between Model 1 and Model 2, but also increases model complexity and makes interpretation of parameters less intuitive. A quadratic model includes three main components, rather than two: initial status, instantaneous rate of change, and curvature. The initial status parameter π_{0i} retains the same interpretation as in the linear growth model. The instantaneous rate of change parameter π_{1i}, however, no longer represents a constant rate of change; it now represents the instantaneous rate of change at a specific point of time (age two in the present case). In fact, the growth rate in quadratic growth model is no longer constant – the rate of change itself now changes smoothly over time as a function of the curvature parameter π_{2i} (Singer and Willett 2003).

One of the most significant changes in coefficients between Models 1 and 2 is that the coefficient for the interaction term between country and sex on height at age two has changed from a nonsignificant value of -0.23 centimeters ($p > 0.05$) to a statistically significant value of 1.80 centimeters ($p < 0.05$). The male advantage in height at age two in the Philippines has also diminished to non-significance. Thus, Chinese boys now show a greater advantage in height at age two – compared to their female counterparts – than is the case for Filipino boys. The instantaneous rate of change equation shows that the male advantage in the growth rate at age two is still greater in China than in the Philipines, though an increase in the standard error means that it drops below conventional levels of statistical significance ($\beta = 0.13$, $p > 0.05$). The curvature equation, however, indicates a declining pattern in the rate of change for the excess male advantage in China, indicating that this Chinese excess male advantage will decline as

variation in both initial status and linear rate of change do exist, even after controlling for all the covariates.

children age. It is worth noting that even though complete or partical “catch-up” growth in height later in life is possible (Golden 1994), substandard physical growth at infancy and early childhood on female children’s physical, psychological, and mental development could have negative consequences for performance in school and other outcomes, regardless of whether catch-up growth occurs.

Model 3 extends Model 2 by adding fixed parameters for three control variables, urban place of residence, mother’s education in years, and mother’s height in centimeters, to assess the extent to which the “difference-in-difference” relationship is explained by differences in these factors across country samples. Comparing Model 3 with Model 2, there are minor changes in coefficients but no change in the substantive findings. For example, these covariates help to explain a small amount of the added advantage experienced by Chinese boys in their initial height (at age two) when compared to Filipino children ($\beta = 1.66$ in Model 3 versus $\beta = 1.80$ in Model 2), but the change is minor. The effects of these three added covariates are in keeping with expectation; children with taller mothers tend to be taller at age two ($\beta = 0.19$) and tend to grow faster at that age ($\beta = 0.04$). One additional year of mother’s education is associated with a 0.17 centimeter increase in height at age two, and a 0.06 centimeter increase in the linear growth rate. Finally, urban children are significantly taller ($\beta = 0.25$) than rural children at age two, and grow significantly faster ($\beta = 0.15$) than rural children (in the instantaneous growth rate). None of the added covariates has a substantial impact on the quadratic growth portion of the model (curvature), despite their statistical significance levels. It should be noted that the sample size for Model 3 is significantly smaller than that for Models 1 and 2 (41,044 versus 50,682 observations) due to missing values in the three newly-added covariates. To assess the influence of missing data on parameter estimates, we re-estimated Model 3 using the multiply imputed data sets described above. Results based on imputed data are reported in Model 4. Estimates of most model parameters differ only slightly between Models 3 and 4, providing some evidence that model results are robust to missing data on covariates.

To demonstrate the predicted growth patterns in height more clearly, we show predicted growth trajectories for boys and girls in China and the Philippines based on Model 2.[11] Figure 4 shows that boys are taller than girls initially, but the gender difference changes as children age. Predicted heights of Filipino girls noticeably exceed those of Filipino boys between ages 4 and 10 years, after which boys catch up and pass girls. This reflects the general finding in the literature that girls experience a growth spurt earlier than boys, though our predicted values show the timing of the growth spurt earlier than typically observed. By contrast, the predicted height of Chinese girls always falls below or just equals the predicted height of Chinese boys, a departure from the typical pattern of sex differences. The unexpectedly early timing of the growth spurt for girls in Figure 4 is an artifact of the quadratic model specification, and we do not interpret these results as indicating the exact ages when girls' average heights are likely to exceed boys'. Rather, this model is a parsimonious, parametric model of reality that summarizes the general patterning of growth trajectories. The key substantive finding, we argue, is that Filipino youth show the expected "crossover" in heights while Chinese youth do not. Thus the results in Table 3 and Figure 4 provide support for our difference-in-difference hypothesis that the male advantage in physical growth is greater in China than in the Philippines.

[11] We base the figure on the predicted values obtained from Model 2 for parsimony, but a figure generated on the basis of Model 3 shows no substantive differences.

Are Findings Similar for Urban and Rural Children?

Thus far, we have controlled for rural versus urban place of residence because prior literature has suggested that the consequences of son preference may be stronger in rural areas (Das Gupta et al. 2003; Marcoux 2002). Now we turn to a more direct comparison of possible urban-rural differences in gender inequality in physical growth.[12] Models 5 and 6 in Table 6 are estimated only for rural children from China and the Philippines and Models 7 and 8 are estimated only for urban children. Models 5 and 7 are estimated using the analytic sample that applies listwise deletion for cases with missing data on explanatory variables, while Models 6 and 8 are estimated using multiply imputed data sets. Only three of the estimated parameters display a notable urban-rural difference: the sex difference in initial height status, the intercept in the quadratic growth rate, and the interaction between sex and country in the linear growth rate. Model 5 shows that there is a significant male advantage in height at age two among rural Chinese children (CHINA × MALE, $\beta = 1.84$, $p < .001$) but not among rural Filipino children (MALE, $\beta = 0.01$), and this effect is substantively the same in Model 6 using imputed data. Model 7 shows a male advantage at age two among urban Filipino children ($\beta = 0.34$) that is not present for rural children (Model 5, $\beta = 0.01$), but the male advantage is not significant using imputed data in Model 8 ($\beta = 0.17$), so we do not have complete confidence in this rural – urban difference. Results for initial height status for Chinese children are substantively the same in

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[12] We also looked at several models that included interaction terms between urban-rural residence and other covariates in the model. However, the models reported in Table 3 are already complicated, and testing these new interaction terms involves a number of three-way even four-way interaction terms, which creates multicollinearity problems and makes the estimates unstable.

urban areas as in rural areas. The urban-rural difference in the intercept in the quadratic growth portion of the model simply means that the growth trajectory for rural children decreases more rapidly than that for urban children, even though rural children have a higher linear growth rate than urban children.

The most interesting urban-rural difference is in the interaction between sex and country in the linear, or instantaneous, rate of growth. The excess male advantage in the growth rate in height in China was estimated to be about 0.13 - 0.17 centimeters at age two in the overall sample, as shown in various models in Table 5. When we stratify the sample by place of residence, the excess advantage for males in China is estimated to be 0.30 centimeters in rural areas and -0.08 centimeters in urban areas, with only the rural figure attaining statistical significance. These results show partial support for our second research hypothesis, that gender discrimination in favor of boys has a greater impact in rural areas of China. Specifically, we find no rural-urban difference in the excess male advantage in children's measured heights at age two, but do find that the excess male advantage in children's linear rate of growth at age two is present in rural areas only. This finding suggests that the process of growth among rural Chinese children is distinctive after early childhood, so that male children's growth trajectories do not fall behind those of females during the typical female growth spurt in the way they do in low son preference settings like the Philippines. Eventually, however, the male advantage diminishes as children reach early adulthood. The potential consequences of this difference in growth patterning for rural Chinese children during adolescence should be explored in future research.

To make rural-urban variation in the gender difference in children's growth more evident, Figures 5 and 6 are generated from Models 5 and 7, respectively. Figure 5 plots the predicted male advantage in height by age and country in rural areas, while Figure 6 plots the same information for children in urban areas. Comparison of Figures 5 and 6 reveals that Chinese male children's height advantage as they age is nearly always greater in rural than in urban areas; males always exceed or equal females in height (male advantage is never less than 0 in Figure 5). The predicted gender difference in height looks much more similar for Chinese and Filipino children in urban areas than in rural areas, with females in both countries showing the expected height advantage over males for a brief period (though the period is briefer and the advantage smaller in China). As for Figure 4, we do not argue that Figures 5 and 6 denote the specific ages at which female or male children's growth spurts occur, or the specific dimension of the gender difference. Nonetheless, the figures demonstrate what models in this analysis and existing theory would predict: that the consequences of son preference are likely to be more notable in rural than in urban areas.

Sensitivity Analyses

The final step in the analysis is to explore the possibility that sibship composition or province- or ethnicity-based differences alter our conclusions about the excess advantage in growth trajectories identified for Chinese males. Table 5 displays the AIC and BIC fit statistics as well as the point estimates of the gender difference in z scores for these models predicting Chinese children's height-for-age z scores; Model 9 represents a baseline model with only the child's age, sex, and urban-rural residence as covariates, and remaining models are compared with Model 9 to assess improvements in model fit and changes in the estimated size of the male advantage.[13]

[13] Models 18 through 20 may not be directly compared to Model 9 because missing data on the ethnicity group of membership reduces the size of the analytic sample in the later models. Models 18 and 19 can be compared, however, in the same fashion as other, similar pairs (such as Models 10 and 11).

One conclusion to be drawn from Table 5 is that many of these explanatory factors improves model fit in an additive sense; specifically, the total number of siblings, province of residence, and ethnic group of membership improve our ability to predict the average child's growth pattern, according to comparisons of AIC and BIC between Model 9 and Models 10, 16, 18, and 20. However, none of the models with interaction terms show evidence for gender-specific effects of the additional explanatory variable(s): this is evident in the comparison of fit statistics for Model 11 versus 10, Model 13 versus 12, Model 15 versus 14, Model 17 versus 16, and Model 19 versus 18. For example, comparing Model 16 against Model 9 shows that distinguishing province of residence improves model fit considerably, but adding the interaction between province and sex in Model 17 leads to a deterioration in model fit (compared to Model 16). These results suggest that the excess Chinese male advantage in height does not vary significantly across provinces of China and is not a function of the ethnic composition of this sample or differences in sibship composition across households, lending support to the robustness of our findings.The last column of Table 5 reports the coefficient associated with the child's sex along with the associated standard errors and significance levels. It is clear that the point estimates for the male advantage in height-for-age remains relatively stable from model to model (0.05~0.10), especially for those models without sex interaction terms (0.05~0.08). While levels of statisitical significance for the male coefficient vary between models as standard errors change, we conclude that the excess Chinese male advantage in growth in height identified above is not likely to be explained by differences across households in sibship composition, province of residence, or ethnic identity. The results shown in Table 5 cannot be directly compared to those in Tables 3 and 4, but provide a useful check on our main results with an alternate metric for height and modeling strategy.

Discussion

This study was designed to explore potential consequences of gender discrimination for the well-being of surviving female children by isolating an impact of son preference on children's growth in height over early childhood and adolescence. We compared China, a societal context 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. This study represents an advance in its use of high quality longitudinal measurements of height and multilevel growth models, and controls for key household characteristics including urban-rural place of residence, mother's height and mother's education. We find support for our first hypothesis; specifically, the male advantage in children's growth trajectories is larger in China than in the Philippines. Our second hypothesis stated that this excess male advantage would be larger for rural Chinese children than for their urban counterparts; we find that at age two males show the same level of excess advantage in height in urban and rural settings, counter to our hypothesis. However, the male excess advantage in the rate of growth at age two is significant only in rural areas of China, suggesting that patterns of growth in height, and gender differences therein, may diverge in rural and urban areas after early childhood.

Our main result – evidence of a male advantage in children's growth in China – counters findings in the recent empirical literature. Other recent studies using a wide range of cross-sectional samples of children under the age of five have shown no male advantage in children's height-for-age z scores (Marcoux 2002) and no disadvantage in female household members' access to calories (DeRose, Das and Millman 2000). One potential reason that we find a male advantage in this Chinese sample is our use of longitudinal data that cover a much wider age range – from birth through adolescence – whereas other studies are limited to samples of young children. We are also able to apply multilevel models of growth trajectories for individual children, rather than looking at relative height-for-age z scores at a single point in time. This allows effective control for individual hegerogeneity and makes the separation of with-individual and between-individual variation possible. These are conventional reasons why results from longitudinal studies are preferred over those from cross-sectional studies, and may allow us to isolate a male advantage in a high son preference context that has been difficult to find in prior studies. Further support for the robustness of these findings can be drawn from the similarity of conclusions drawn from results obtained using raw height measurements (Tables 3 and 4) and those using *z* scores (Table 5), the more commonly used metric in studies of child stature.

More in keeping with existing theory and empirical findings, our results suggest that the negative effects of son preference on girl children appear to operate more strongly in rural areas of China, at least in terms of growth rates in early childhood. For example, Clark (2000) reports that in India, rural women desire a greater 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 only present among rural Chinese women, not urban women (Marcoux 2002). The present study may be the first, however, 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 across rural and urban areas. Nonetheless, comparisons of Chinese and Filipino children's achieved heights at age two showed greater excess advantage for Chinese males regardless of urban or rural place of residence, suggesting that differential treatment of male and female children may be widespread in China.

Previous studies have shown that a child's birth order and the number, sex and ordering of her siblings matters for her survival and well-being (Clark 2000; Pande 2003). Our results suggest that the number of siblings has important implications for childrens' growth (Model 9 in Table 5), but does not explain the excess male advantage in China. Nonetheless, this finding suggests that future analyses focusing more carefully on sibship composition and childrens' growth trajectories across different societal contexts are warranted. Similarly, province of residence and ethnic group identity help explain children's growth trajectories (Table 5), but not the male excess advantage in China that we attribute to strong preference for sons.

The choice to directly compare the growth of children between societies with very different levels of son preference, from very high (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 of discrimination against girl children on their growth and well-being. First, because of the prevalence of sex-selective abortion and infanticide in China (Banister 2004; Ren 1995), our Chinese sample will automatically exclude those girls would have been most disadvantaged, had they survived gestation and early life. Parents who are least interested in having girl children and would be most likely to discriminate against them may have opted out of having daughters altogether (Goodkind 1996). Second, survey attrition due to child mortality in the samples used here, to the extent that it has occurred, would likely remove from the study those children in poorest health and potentially the female children experiencing the greatest levels of discrimination. Thus, our estimates of gender difference in children's growth in height in China may fall toward the lower bound of the true association between the level of son preference in a given context and gender disparities in children's growth.

Some limitations of the present study should be considered when interpreting these results. Perhaps most importantly, we lack data on parents' behavior toward male and female children, the mechanism purported to underlie our findings. We instead compare children's growth patterns to provide objective "evidence" of 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 male growth advantage among children and adolescents. In addition, different survey designs mean that the sample of Chinese children is dominated by those living in rural areas, where son preference is expected to have the strongest effects, while the Filipino sample has a relatively small number of children 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 ample. 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 (Marcoux 2002; Obermeyer and Cardenas 1997). The present study, however, suggests that cultural preferences for sons do have consequences for the health and well-being of females throughout the early years of the life course. 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 parent's behavior and have consequences for demographic outcomes. Taken together, such findings suggest that intervention to improve the conditions of girl children could have important consequences over the entire life course. We have shown that during the years crucial for eduational 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 literatures of demography and social stratification; in these literatures, one set of studies has shown that son preference influences female infant survival and early life health and treatment (Graham, Larsen and Xu 1998; Short et al. 2001). Another set of studies has linked height deficits in early childhood to serious negative implications for school performance and labor market position in later life (Daniels and Adair 2004; Glewwe, Jacoby and King 2001; Jamison 1986). The present study has bridged the life course gap between early childhood and later adolescence, to show that conditions of high son preference may generate gender-differentiated growth trajectories, particularly disadvantageous for rural girls, 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.

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Table 1. Cross-Wave Comparison of Selected Variables in the Cebu Longitudinal Health and Nutrition Survey

	Height (cm.)	Age (years)	% Male	% Urban	Mother's Education (years)	Mother's Height (cm.)
At Birth	49.33	0.0	53	76	7.70	150.83
2nd Month	56.38	0.16	53	76	7.66	150.85
4th Month	61.11	0.33	53	75	7.64	150.83
6th Month	64.34	0.49	53	75	7.61	150.83
8th Month	66.86	0.66	53	75	7.60	150.83
10th Month	69.01	0.82	53	75	7.60	150.82
12th Month	70.81	0.99	53	75	7.60	150.86
14th Month	72.40	1.16	53	75	7.58	150.87
16th Month	73.81	1.32	53	75	7.58	150.82
18th Month	75.21	1.49	53	75	7.59	150.85
20th Month	76.57	1.66	53	75	7.58	150.84
22nd Month	77.81	1.82	53	75	7.57	150.86
24th Month	79.18	1.99	53	75	7.54	150.84
1991 Follow-up	117.70	8.49	53	74	7.54	150.90
1994 Follow-up	133.78	11.52	52	73	7.67	150.89
1998 Follow-up	154.01	15.50	52	72	7.66	150.92

Table 2. Cross-Wave Comparison of Selected Variables in the Chinese Health and Nutrition Survey

	Height (cm.)	Age	% Male	% Urban	Mother's Education (years)	Mother's Height (cm.)	Standardized Height-for-Age *z* Scores
1989	89.79	2.71	53	30	6.85	155.43	-1.32
1991	100.55	4.12	53	29	6.75	155.32	-1.37
1993	109.76	5.53	53	29	6.72	155.36	-1.25
1997	125.50	8.16	53	29	6.98	155.69	-0.99
2000	134.76	9.75	53	29	7.37	156.04	-0.84

Table 3. Individual Growth Models of Children's Height using Pooled Chinese CHNS and Filipino CLHNS Samples

	Model 1	Model 2	Model 3	Model 4 Mult. Imp.
FIXED EFFECTS				
Initial Status				
Intercept	73.46***	76.36***	46.47***	47.12***
	(0.09)	(0.10)	(1.47)	(1.37)
Male	1.09***	0.14	0.27	0.11
	(0.13)	(0.13)	(0.14)	(0.12)
China	9.41***	3.74***	2.83***	2.91***
	(0.17)	(0.18)	(0.19)	(0.18)
China × Male	-0.23	1.80***	1.66***	1.76***
	(0.23)	(0.24)	(0.24)	(0.23)
Urban Residence			0.25*	0.30*
			(0.12)	(0.11)
Mother's Education			0.17***	0.11***
			(0.01)	(0.01)
Mother's Height			0.19***	0.19***
			(0.01)	(0.01)
Rate of Change (Linear)				
Intercept	6.38***	8.90***	2.83***	3.32***
	(0.02)	(0.03)	(0.45)	(0.43)
Male	-0.11***	-0.83***	-0.88***	-0.85***
	(0.02)	(0.04)	(0.04)	(0.04)
China	-0.31***	-0.93***	-0.77***	-.95***
	(0.03)	(0.06)	(0.06)	(0.06)
China × Male	0.15***	0.13	0.17*	0.16*
	(0.04)	(0.08)	(0.08)	(0.07)
Urban Residence			0.15***	0.20***
			(0.04)	(0.03)
Mother's Education			0.06***	0.06***
			(0.00)	(0.00)
Mother's Height			0.04***	0.03***
			(0.01)	(0.00)
Rate of Change (Quadratic)				
Intercept		-0.27***	-0.10*	-0.12*
		(0.00)	(0.04)	(0.04)
Male		0.09***	0.09***	0.09***
		(0.00)	(0.00)	(0.00)
China		0.11***	0.08***	0.10***
		(0.00)	(0.00)	(0.00)
China × Male		-0.03***	-0.03***	-0.03***
		(0.01)	(0.01)	(0.01)
Urban Residence			-0.01**	-0.01**
			(0.00)	(0.00)
Mother's Education			-0.00***	-0.00***
			(0.00)	(0.00)
Mother's Height			-0.00**	-0.00**

			(0.00)	(0.00)
VARIANCE COMPONENTS				
Residual	5.33***	4.37***	4.37***	4.37***
	(0.02)	(0.02)	(0.02)	(0.00)
Initial Status	3.06***	3.30***	2.84***	2.90***
	(0.05)	(0.05)	(0.05)	(0.02)
Rate of Change (Linear)	0.47***	0.43***	0.36***	0.37***
	(0.01)	(0.01)	(0.01)	(0.02)
Log Likelihood	-162464	-154032	-124065	NA
D. F.	11	15	24	24
AIC	325314	308094	248178	NA
BIC	325411	308226	248385	NA
N	50,682	50,682	41,044	50,682

Note: Numbers in parentheses are standard errors. Model 3 and 4 use the same covariates but different analytic samples. Model 3 is estimated on listwise deleted data, while Model 4 is estimated on multiply imputed data (number of multiply-imputed data sets = 5). Likelihood-based statistics are not reported for Model 4 because multiple imputation does not involve calculating the likelihood, and applying Rubin's rule to these parameters does not work reliably. * $p < .05$; ** $p < .01$; *** $p < .001$

Table 4. Quadratic Growth Models of Children's Height by Urban versus Rural Place of Residence, CHNS and CLHNS samples.

	Rural		Urban	
	Model 5	Model 6 Mult. Imp.	Model 7	Model 8 Mult. Imp.
FIXED EFFECTS				
Initial Status				
Intercept	46.00***	48.56***	45.68***	45.08***
	(2.49)	(2.32)	(1.83)	(1.66)
Male	0.01	-0.06	0.34*	0.17
	(0.30)	(0.26)	(0.15)	(0.13)
China	2.59***	2.70***	2.80***	2.97***
	(0.29)	(0.27)	(0.30)	(0.28)
China × Male	1.84***	1.89***	1.78***	1.83***
	(0.39)	(0.35)	(0.41)	(0.39)
Mother's Education	0.15***	0.10***	0.20***	0.12***
	(0.03)	(0.03)	(0.02)	(0.01)
Mother's Height	0.20***	0.18***	0.19***	0.20***
	(0.02)	(0.02)	(0.01)	(0.01)
Rate of Change (Linear)				
Intercept	4.44***	5.00***	1.90***	2.30***
	(0.75)	(0.72)	(0.58)	(0.54)
Male	-0.92***	-0.91***	-0.86***	-0.82***
	(0.08)	(0.07)	(0.05)	(0.04)
China	-0.90***	-1.06***	-0.57***	-0.83***
	(0.09)	(0.08)	(0.10)	(0.10)
China × Male	0.30**	0.31**	-0.08	-0.07
	(0.11)	(0.11)	(0.13)	(0.13)
Mother's Education	0.05***	0.06***	0.06***	0.06***
	(0.01)	(0.01)	(0.01)	(0.01)
Mother's Height	0.03***	0.02***	0.04***	0.04***
	(0.00)	(0.00)	(0.00)	(0.00)
Rate of Change (Quadratic)				
Intercept	-0.23***	-0.26***	-0.01	-0.03
	(0.06)	(0.06)	(0.05)	(0.05)
Male	0.09***	0.09***	0.09***	0.09***
	(0.01)	(0.01)	(0.00)	(0.00)
China	0.09***	0.10***	0.07***	0.09***
	(0.01)	(0.01)	(0.01)	(0.01)
China × Male	-0.04***	-0.04***	-0.01	-0.01
	(0.01)	(.01)	(.01)	(.01)
Mother's Education	-0.00***	-0.00***	-0.00***	-0.00***
	(0.00)	(0.00)	(0.00)	(0.00)
Mother's Height	-0.00	0.00	-0.00***	-0.00***
	(0.00)	(0.00)	(0.00)	(0.00)
VARIANCE COMPONENTS				
Residual	4.34***	4.33***	4.37***	4.37***
	(0.03)	(0.01)	(0.02)	(0.00)
Initial Status	3.16***	3.20***	2.53***	2.70***
	(0.09)	(0.03)	(0.06)	(0.02)

Rate of Change	0.39***	0.40***	0.34***	0.35***
	(0.01)	(0.03)	(0.01)	(0.03)
Log Likelihood	-48706	NA	-75324.4	NA
D. F.	21	21	21	21
AIC	97455	NA	150690.7	NA
BIC	97616	NA	150861.5	NA
N	15,922	18,326	25,122	32,353

Note: Numbers in parentheses are standard errors. Models 5 and 6 (and Models 7 and 8) use the same covariates but different analytic samples. Models 5 and 7 are estimated on listwise deleted data, while Models 6 and 8 are estimated on multiply imputed data (number of multiply-imputed data sets = 5). Likelihood-based statistics are not reported for Models 6 and 8 because multiple imputation does not involve calculating the likelihood, and applying Rubin's rule to these parameters does not work reliably.

* $p < .05$; ** $p < .01$; *** $p < .001$

Table 5. Assessing Additive and Interactive Effects of Sibship Composition, Province of Residence, and Ethnic Minority Group Membership on the Gender Difference in Standardized Height-for-Age *Z* Scores, Chinese CHNS sample.

Model	Number of Cases	Deg. of Freedom	*AIC*	*BIC*	Male Coefficient
M 9: Baseline Model	11,460	7	34086.8	34138.2	0.07* (0.03)
M 10: M 9 + # of Siblings	11,460	10	33945.2	34018.7	0.05 (0.03)
M 11: M 10 + # of Siblings × Male	11,460	13	33943.7	34039.2	0.06 (0.05)
M 12: M 9 + Elder Brother/Sister	11,460	9	34082.9	34149.1	0.06 (0.03)
M 13: M 12 + Elder Brother/Sister × Male	11,460	11	34086.7	34167.5	0.07 (0.04)
M 14: M 9 + Birth Order	11,460	10	34085.0	34158.4	0.08* (0.03)
M 15: M 12 + Birth Order × Male	11,460	13	34090.8	34186.3	0.08* (0.04)
M 16: M 9 + Province	11,460	15	33680.0	33790.2	0.07* (0.03)
M 17: M 16 + Province × Male	11,460	23	33686.7	33855.7	0.09 (0.10)
M 18: M 9 + Ethnicity	9,609	8	27876.6	27934.0	0.06 (0.03)
M 19: M 18 + Ethnicity × Male	9,609	9	27878.4	27943.0	0.10 (0.09)
M 20: M 9 + Province + Ethnicity	9,609	16	27590.1	27704.8	0.06* (0.03)

Note: Model 9 includes controls for child's age, sex, urban/rural place of residence; Models 10 through 20 add indicators of sibship composition, province of residence, or ethnic minority group membership and associated interactions with child sex. See text for descriptions of specific models. The last column reports point estimates for the male coefficient along with associated standard errors and significance levels;
* $p < .05$; ** $p < .01$; *** $p < .001$

Figure 1. Distribution of Filipino Children's Height by Age

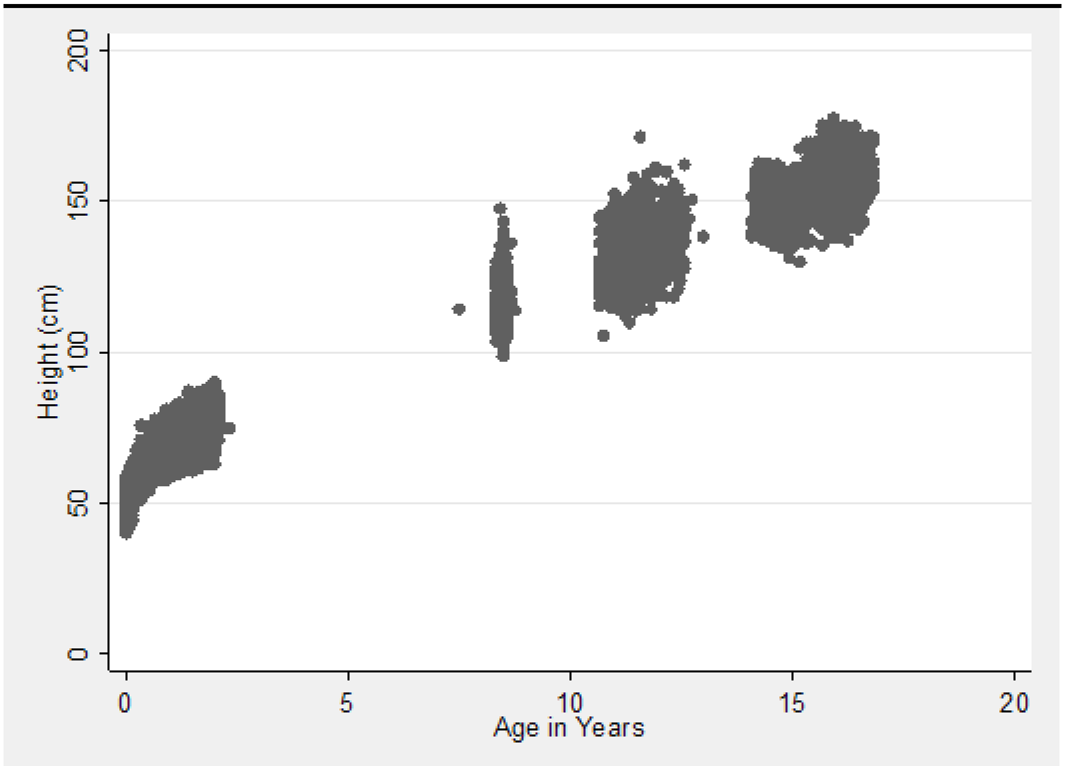


Figure 2. Distribution of Chinese Children's Height by Age

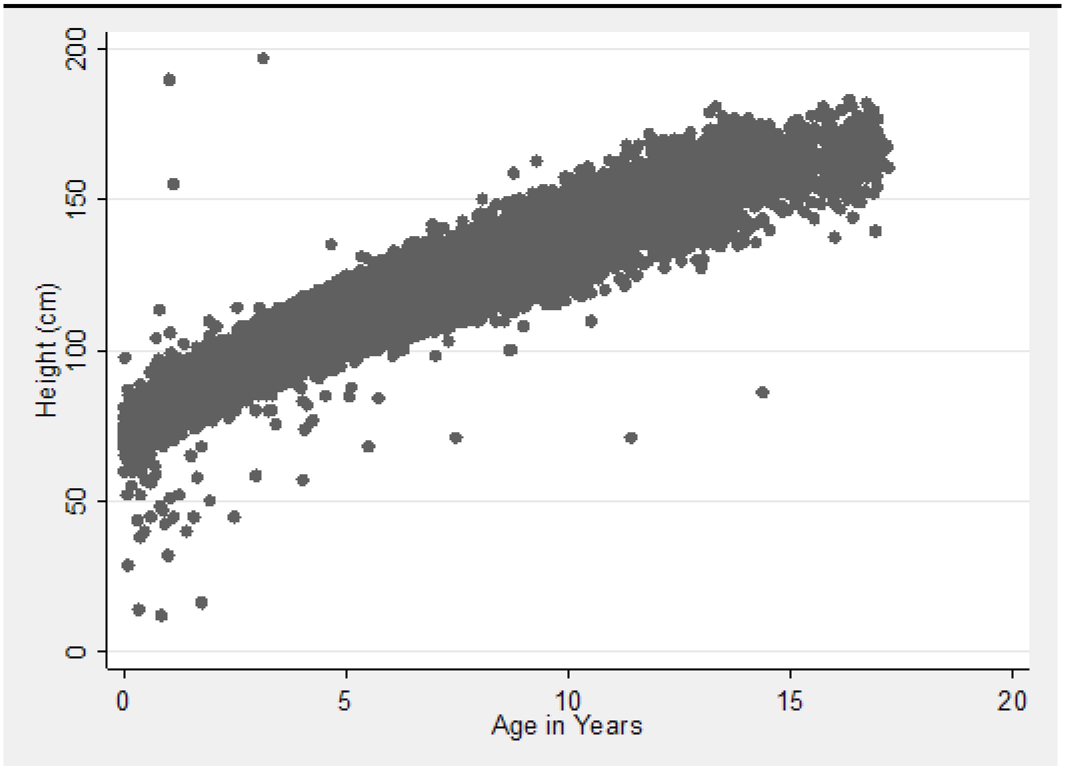


Figure 3. Smoothed Age-Specific Male Advantage in Mean Height by Country

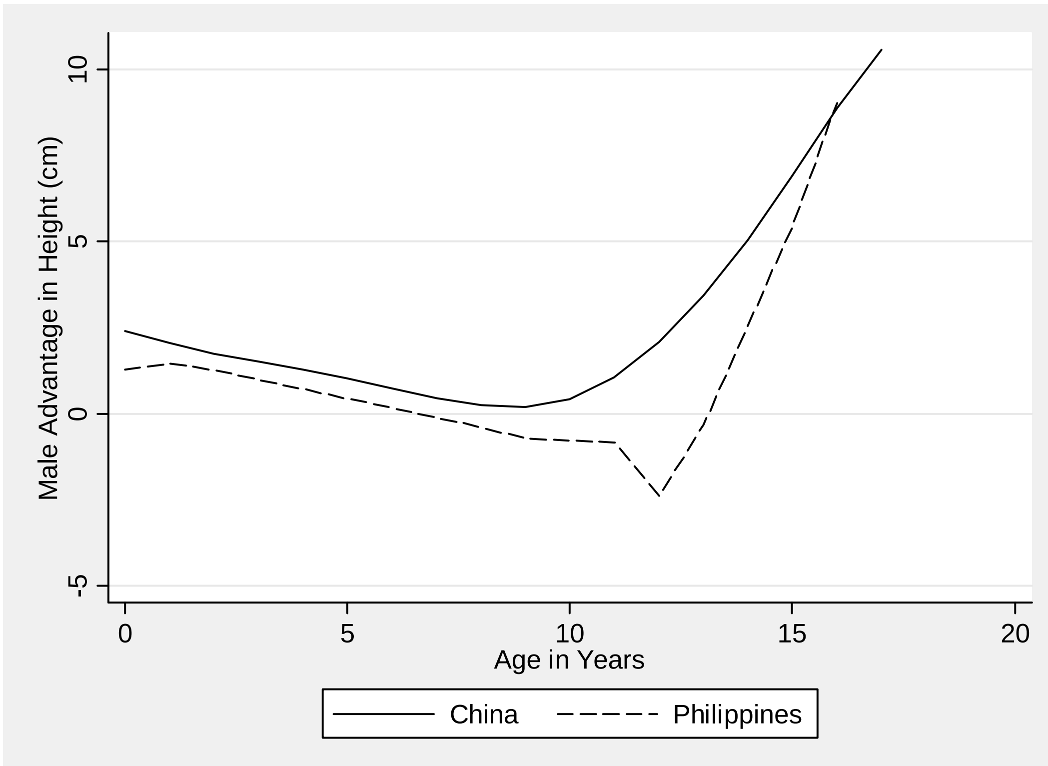


Figure 4. Predicted Children's Growth Trajectories by Country and Sex

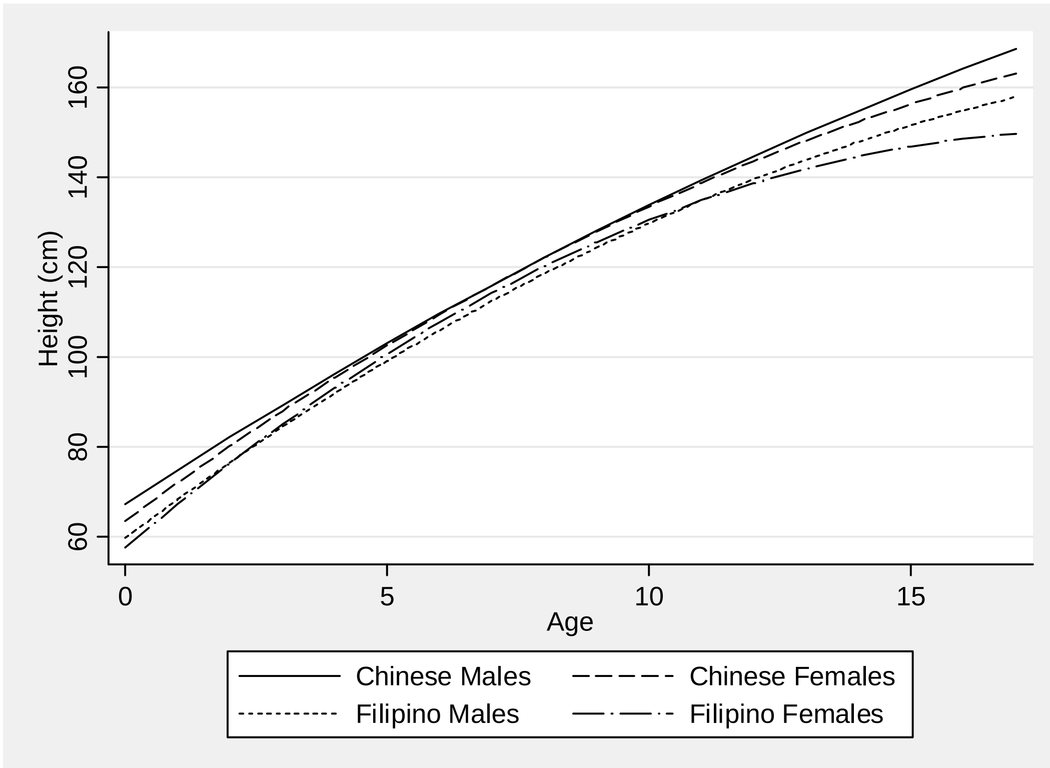


Figure 5. Predicted Age-Specific Gender Difference in Height by Country, Rural Areas

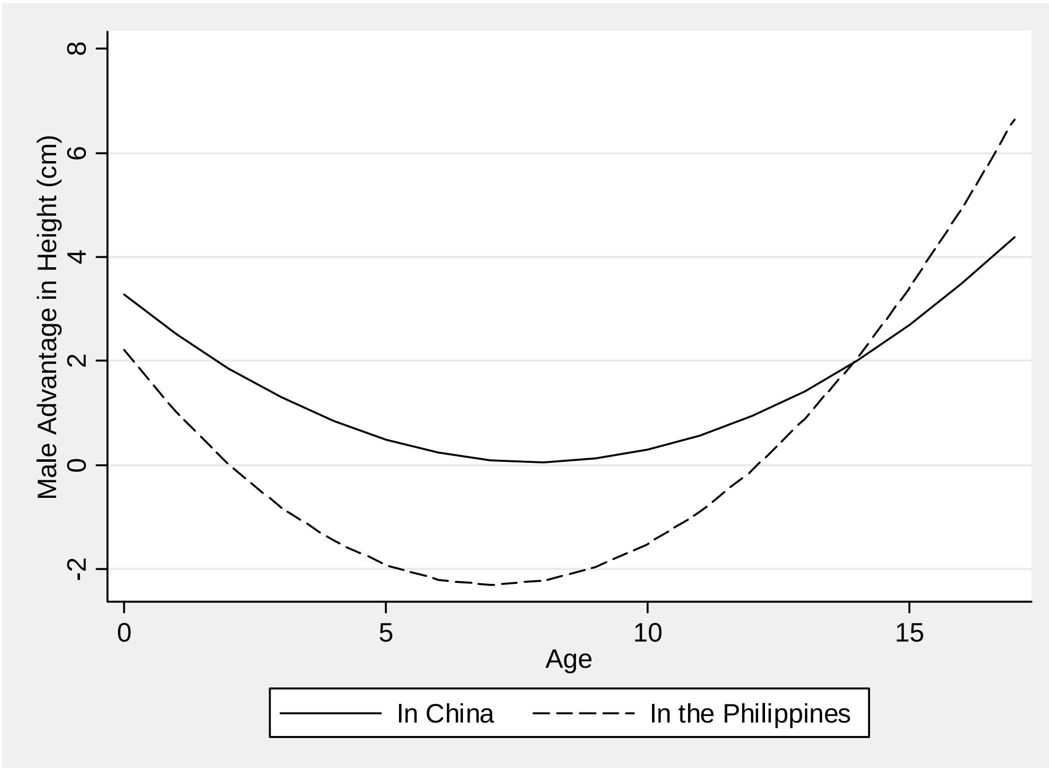
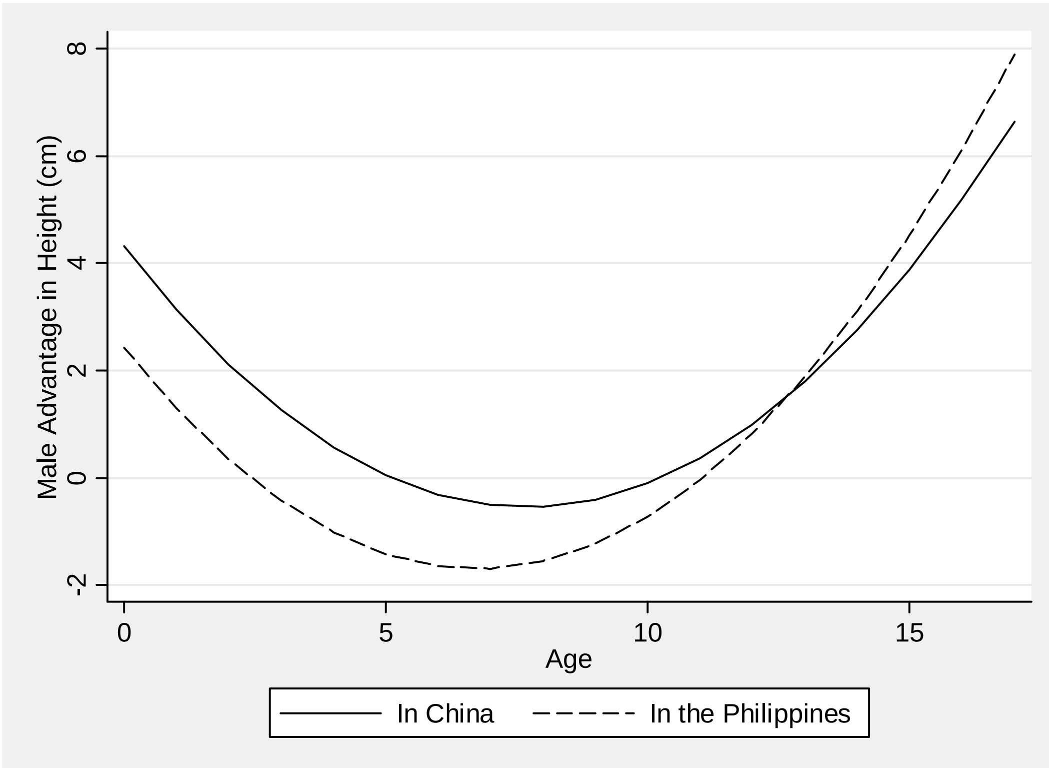


Figure 6. Predicted Age-Specific Gender Difference in Height by Country, Urban Areas



Appendix A: Sex- and Age-Specific Summary Statistics for Children's Height from the Cebu Longitudinal Health and Nutrition Survey

Age	Female			Male		
	Mean	S.D.	N	Mean	S.D.	N
0	55.16	5.35	4,177	56.16	5.75	4,748
1	68.81	4.29	7,062	70.34	4.33	7,856
2	76.34	3.81	4,200	77.98	3.71	4,642
3	71.67	2.08	3	--	--	0
8	118.07	5.86	231	117.89	5.27	272
9	117.58	5.59	818	117.68	5.63	900
10	112.20	--	1	--	--	0
11	133.33	7.33	463	130.50	6.56	518
12	136.83	7.58	571	133.84	6.95	619
13	140.69	8.27	13	136.89	11.30	11
14	148.83	5.44	117	--	--	0
15	149.02	5.56	791	156.26	7.10	43
16	150.49	5.59	64	158.49	6.74	884
17	--	--	0	159.23	6.65	132

Appendix B: Sex- and Age-Specific Summary Statistics for Children's Height from the Chinese Health and Nutrition Survey

Age	Female			Male		
	Mean	S.D.	N	Mean	S.D.	N
0	72.64	7.68	157	74.93	6.95	177
1	80.64	8.67	373	82.62	8.87	443
2	88.46	7.44	401	90.48	5.62	461
3	96.08	5.41	446	97.58	7.23	530
4	102.73	6.10	460	103.27	6.37	510
5	108.86	5.82	445	109.68	6.39	539
6	114.31	6.65	427	115.88	6.82	459
7	120.39	7.21	413	121.09	6.83	457
8	125.93	7.21	342	126.37	7.02	408
9	132.20	8.18	323	131.80	8.04	380
10	137.14	8.83	330	136.06	8.23	341
11	143.75	9.90	247	143.06	8.33	303
12	149.66	7.74	231	149.63	9.60	253
13	152.86	6.81	232	154.59	10.43	250
14	155.00	8.30	178	160.74	8.88	199
15	156.77	6.00	78	164.21	6.87	79
16	157.46	6.60	69	168.34	7.35	73
17	160.23	6.20	35	169.01	6.19	40