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A Story of Increased Gender Inequality?**

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Child Nutrition in India in the Nineties: A Story of Increased Gender Inequality?

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Abstract

We establish some new interesting stylized facts on the changes in boy versus girl nutritional status in India during the nineties, a period of rapid economic growth. Our analysis is based on the comparison, over time and across genders, of the distribution of z-scores calculated for height and weight measures. Overall, we find that child nutrition improved substantially, but we also find that gender differences in nutritional status increased as well, with nutritional status improving substantially more for boys than for girls. Consistent with a large literature that shows the existence of a steep North-South gradient in gender inequality in India, we find that changes in nutritional status appear to be much more similar between genders in the South. We also estimate predicted changes in nutritional status based on changes in the distribution of household wealth (proxied by asset ownership) and a few other observed household characteristics. Actual changes appear to be relatively close to predicted ones in urban areas. For children living in the rural sector the results are more mixed, and we observe that actual changes in weight are quite larger than predicted ones for boys, while they are much worse than the predicted ones for girl height. We also estimate that the predicted changes are generally larger for boys than for girls.

JEL: I12, J13, O53

Key words: Child Nutrition, India, Child Anthropometry

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1 Introduction

The main goal of this paper is to establish some stylized facts on the changes in boy versus girl nutritional status during the nineties, and to attempt to identify how much of these changes can be explained by changes in household wealth and other observable household characteristics. Overall, we find that child nutrition improved substantially, but that gender inequality in nutritional status increased as well, with nutritional status improving substantially more for boys than for girls. We also document the existence of apparent geographical differences in these changes.

Gender inequality is a well-known and still widespread reality in India, particularly in the North. One of its most noted manifestation is the unnaturally high ratio of men to women in these areas (see, among the countless others, Basu (1992), Dasgupta (1993), Miller (1981), Murthi, Guio, and Drèze (1995) and references therein). Previous research suggests that among its proximate pathways are sex differences in domains such as abortion, infanticide, child health care, and child nutrition. Several studies from such different disciplines as Anthropology, Economics and Sociology have found that son preference is particularly strong in areas where the cultural, social, and economic role of women in society and/or within the household is weaker. So, for example, excess female mortality among young girls is more common in areas where the role of women as bread earners is smaller, dowries are more common, or bequests tend to favor sons over daughters.

Many of the factors associated with son preference appear to have economic content.¹ For this reason, economists have explored the possibility that household utility maximization might entail a certain degree of preferential treatment for boys, even if the welfare of boys and girls enter equally into the parents' utility function. Son preference, in this sense, might therefore be an unfortunate but rational response to unequal economic "returns" to boys and girls. In a seminal paper, Rosenzweig and Schultz (1982) use census and survey data to argue that parents allocate more resources to boys—therefore improving their survival rates relative to girls—when employment opportunities become more biased in favor of men. Behrman (1988a) and Behrman (1988b), using data from a small number of Indian villages, find that parents favor equal treatment of children, but also find evidence of pro-male bias in intrahousehold allocation of resources during the *lean* season, when resource constraints are more likely to bind. Jensen (2002), building on insights from Yamaguchi (1989), discusses how gender bias in average outcomes may also arise without any unequal treatment, if girls are more likely to live in households with more siblings. The existence of this phenomenon has indeed been documented and it has been explained as a consequence of

¹There are clear exceptions, such as the importance of males in performing certain religious rituals, which is especially common in North India.

differential stopping fertility behavior, whereby families are willing to have more children when they have not yet achieved the desired number of sons. From this perspective, girls' *mean* outcomes may not be worse because females are discriminated against in the intrahousehold allocation of resources, but because they are more likely to live in families with less resources per head.

The fact that resource constraints—coupled with pro-male bias in economic opportunities—appear to provide an economic “rationale” for the existence of gender bias, might lead one to expect a path towards equalization as a consequence of economic development, if this is accompanied by an increase in the resources available to households. However, it has been observed that female discrimination in India is *not* limited to the poorest and least educated households. In fact, in some studies it actually appears to be more frequent among certain high castes (Das Gupta (1987)). Similarly, it has been suggested that the decline in fertility that has accompanied economic development in India may have contributed to a *worsening* of gender bias, as the desired number of sons may have decreased less quickly than the desired total number of children (Das Gupta and Bhat (1995), Basu (1999)). Anderson (2003) constructs a model where economic development, in a caste-based society, leads to an *increase* in dowries. This might lead to an increase in son preference. Goldin (1995), among others, documents the existence of a U-shaped female labor force participation rate as a function of economic development, so that the role of women as bread earners might *decrease* in the first stages of development. Overall, these observations lead to ambiguous predictions on the relation between son preference and economic development.

Child weight and height performance can be viewed as the output of a health production function whose inputs include elements such as nutritional intakes, exposure to infections, and health care (as well as, of course, genetic predisposition). In this sense, height and weight are affected by virtually all of the pathways through which gender bias operates. Anthropometric indicators are also extremely important because there is a well documented relationship between child malnutrition and poor adult outcomes (see, for example, Strauss and Thomas (1998)). When evaluating gender differences, another advantage of nutritional status versus, say, nutrient intakes, morbidity, or health care, is that the former is easily measured, and therefore much less prone to measurement error or reporting bias.

India experienced several years of fast economic growth during the nineties, years that according to several observers have also witnessed a large reduction in poverty (see, for example, Deaton (2003), Deaton and Drèze (2002), Tarozzi (2004), or, for a more skeptical view, Datt, Kozel, and Ravallion (2003) or Sen and Himanshu (2004)). The goal of this paper is to document new interesting stylized facts over this period on the change in boy versus girl nutritional status as measured by anthropometric indicators, namely weight-for-age, height-for-age, and weight-for-height. For this

purpose, we use data from the Indian National Family and Health Survey, a data set that contains detailed information on health and fertility for two cross-sections of ever married women of fertility age, the first from 1992-93 and the second from 1998-99.

To evaluate changes in nutritional status, we transform the anthropometric indicators into *z-scores*, that is, we normalize the indicators by using mean and standard deviation of the same index for children of the same age and gender in a reference population. The use of *z-scores* is common in nutritional studies (more on this below), and two reasons make it particularly useful for our purposes. First, it facilitates comparisons between genders, as nutritional status is evaluated *relative* to children of the same gender in a reference population. Second, it allows to pool together children of any age, so that one can simply evaluate the overall nutritional status in a population estimating nonparametrically the whole *distribution* of the *z-scores*. Indeed, this second advantage of using *z-scores* is crucial for our purposes, as most of our results are based on the comparison of *z-score* cumulative distribution functions between genders (for a given wave) or over time (for a given gender). Overall, our findings suggest that child nutrition improved substantially, but the changes over time of the distributions of *z-scores* show a tendency towards male advantage in nutritional status. Consistently with a large literature that show the existence of a steep North-South gradient in gender inequality in India, we find that changes in nutritional status appear to be much more gender neutral in the South.

We also estimate predicted changes in nutritional status based on changes in the distribution of household wealth (proxied by asset ownership) and a few other observed household and individual characteristics. Actual changes appear to be close to predicted ones, especially in urban areas. Interestingly, we also find that the *predicted* changes are generally larger for boys than for girls. Overall, our results provide some support to the view that poverty did decline in India during the nineties, while they are difficult to reconcile with the notion, argued by some researchers, that malnourishment has actually *increased* during this period (see Ray and Lancaster (2005)).

The paper proceeds as follows. In the next section we describe the dataset. In Section 3 we discuss the anthropometric indicators we choose as outcomes for our analysis. In Section 4 we document the extent of gender differences in child nutritional status, and we study how the distribution of anthropometric indices changed between the two NFHS waves. In Section 5 we study how much of the observed changes can be explained by changes in a few selected household and individual characteristics, and Section 6 concludes.

2 Data

The main source of our data is the two waves of the Indian National Family and Health Survey (NFHS) available at the time of writing. The NFHS is one of the many Demographic and Health Surveys that have been carried out in several developing countries—adopting a largely standardized questionnaire—with the primary purpose of collecting information on health, fertility and other family issues from ever married women of fertility age. The first wave (NFHS-I) was carried out between April 1992 and August 1993 on a sample of ever married women of age between 13 and 49. The second wave (NFHS-II) was completed between November 1998 and December 1999, sampling ever married women of age 15-49.² Each survey contains reports from approximately 90,000 women, sampled from all the Indian states using a stratified and clustered survey design. All statistics reported in this paper make use of the sampling weights contained in the survey.

The largest component of the surveys is an individual questionnaire administered to each ever married woman of fertility age in the sample. The questionnaire also includes information on health, contraception and fertility preferences, as well as a complete birth history and very detailed information on the health status of younger children.³ In particular, height and weight were measured for children below age 4 in NFHS-I, and below age 3 in NFHS-II. Because of lack of appropriate measuring tools, height was not measured during fieldwork in the first states covered by NFHS-I. These states, which formed the so-called *Phase I* of the survey, are Andhra Pradesh, West Bengal, Himachal Pradesh, Madhya Pradesh, and Tamil Nadu. In most of the results that follow we are interested in comparing the distribution of different anthropometric outcomes within the same survey, or changes in such distributions between the two waves. To ensure that the population whose outcomes are being compared is the same, we will base most of our results on states and age groups that are represented in both waves. We will refer to the states for which height was recorded in 1992-93 as *Phase II* states. In both waves, a village questionnaire records information on village characteristics, while a household questionnaire contains several household characteristics, including a complete household roster, and individual information on work status, educational attainment, and a few selected health indicators.

Table 1 reports a list of summary statistics calculated using data from the individual question-

²We ignore the difference in lower bound of mothers' age in the waves, as less than 0.5% of women in NFHS-I were 13 or 14 years old.

³NFHS-II also contains questions related to quality of available health care, the woman's empowerment within the household, AIDS awareness, mother's anthropometric indicators, and mother and children's anemia. We do not use such information as it is not available in the first wave. Also, some observers have raised doubts on the reliability of some of these variables (see Irudaya Rayan and James (2004)).

naires administered to women in the two waves of the NFHS. As it is customary in the literature, we analyze the rural and the urban sector separately. All averages and proportions are calculated using the household sampling weights provided within the survey. The statistics for all India include all valid observations, while the regional means only include the major Indian states, excluding Union Territories (which account for less than 5 percent of the population). For several statistics we also present a geographical breakdown following the geo-cultural classification proposed by Sopher (1980). States are then grouped as follows: North includes Delhi, Gujarat, Haryana, Himachal Pradesh, Jammu, Madhya Pradesh, Punjab, Rajasthan and Uttar Pradesh. Assam, Bihar, Orissa and West Bengal form the Eastern region, while the South includes Andhra Pradesh, Karnataka, Kerala, Maharashtra and Tamil Nadu.

Many indicators suggest that important changes are taking place. There is evidence that fertility is declining, both in cities and in rural areas. Average household size declined by about 0.2 persons, and the number of children below age 5 by about 0.1. However, these results should be considered with caution, as some observers, citing evidence from other data sources, have suggested that the number of births in NFHS-II might be underreported (see Irudaya Rayan and James (2004), and references therein). Desired family size is also declining, and the use of contraceptives has become more common.⁴ In urban areas, the proportion of women *not* practicing any form of birth control declined from 52 to 45 percent. In rural areas the proportion declined from 65 to 58 percent. The proportion of women desiring no more than three children increased from 80 to 86 percent in urban areas, and from 65 to 73 in rural areas. The proportion desiring no more than 2 children increased even more, by approximately 11 percentage points, reaching 67 percent in towns, and 46 percent in the countryside. We do not find evidence that the number of desired boys is declining more slowly than the overall number of desired children. On the contrary, the mean proportion of children that are desired to be girls is *higher* in 98-99 than in 92-93, even if the figure is still below one half. For this statistic we also report averages calculated separately for different geographical areas, in order to highlight the well-known existence of strong regional patterns in son preference. Even in the North, the proportion of desired girls *increases* by approximately one percentage points in both rural areas (where it was 38% in NFHS-I) and in towns (where it was 41.9%). In the South the proportion increases from 46.3 to 47.2 percent in cities, and from 43.5 to 45.6 percent in rural areas. Similar patterns emerge in Eastern states.⁵

⁴Desired family size is defined as the ideal number of children that a respondent with no children would like to have, or the ideal number she would have liked to have if she could go back to the time she did not have any children.

⁵If we interpret non-numerical responses—which may include answers such as “up to God”—as expressing indifference with respect to child gender, the results are qualitatively identical, with only a generalized small decrease in son preference, which arises by construction.

Looking at female labor force participation, three patterns are apparent. First, in every region, and in both waves, women are much more likely to work in rural than in urban areas, where participation rates are about 40% lower than in the countryside. Second, participation rates have increased over time in all areas, especially in the North, where participation rates increased from 16.2 to 21.2 percent in urban areas, and from 29.5 to 37.3 in rural areas. Third, participation rates are about twice as large in the South than in the North, both in cities and in villages. For example, in NFHS-II 61.2 percent of women worked in the rural South, while only 37.3 percent did in rural North. In Eastern states women participation is even lower than in the North. Turning to the proportion of working women who do so for a salary, Table 1 shows that this has remained remarkably stable over time, and that in urban areas it remains close to 90 percent in all regions. While in rural South approximately three quarters of working women earn a salary, the proportion is only two thirds as large in the North. These statistics are clearly very rough measures of the role of women in the economy, but overall they do not lend support to the hypothesis that India is along the downward path of the U-shaped curve that frequently characterizes the relation between women's economic role and development (see Goldin (1995)).

Female illiteracy rates once again confirm a familiar North-South pattern. In 1992-93, almost 80 percent of ever married women of fertility age that live in Northern states had no formal education. In the South the proportion was still very high, but 20 percentage points lower. The gradient is also clearly present in urban areas, but at levels approximately 50 percent lower. Illiteracy is significantly less common among the sample women's partners, with a proportion usually half as large than for women. Note also that there is no clear North-South gradient in illiteracy for men, so that one cannot easily interpret the gradient in women's illiteracy as indicating geographical differences in availability of, or attitudes towards, schooling. All these patterns are still present in 1998-99, but there are clear signs of improvements over time, as formal education is becoming more common both for men and for women. The last rows of Table 1 show a remarkable increase in the proportion of both men and women with at least a secondary degree. In urban areas of all regions the percentages for women are approximately three times as large in NFHS-II as in NFHS-I. Overall, the proportion increases from 10.7 to 32.8. The figures are much lower in rural areas, but the increase is even larger, in relative terms, as the overall proportion of women with at least a secondary degree increases from 0.8 to 7.7 percent.

Table 2 shows selected summary statistics for child outcomes. The top part of the table shows that a geographical pattern in son preference also emerges looking at the gender gap in mortality rates. Because each woman was asked to report a complete birth history, we can calculate a measure of male versus female mortality rates. For each interviewed woman who reported giving birth to

at least one male and at least one female, we compute the difference between the proportion of surviving males and that of surviving females.⁶ In line with the existing literature, we find that gender bias is concentrated in the Northern states, especially in rural areas, where in 1992-93 the fraction of surviving boys was approximately 2.5 percentage points higher than the fraction of surviving girls. In this region, there is a slight decline in the differential mortality gap between NFHS-I and NFHS-II. In urban areas the gap is again positive, but extremely small, declining over time, and not statistically different from zero using standard significance levels. We also find a small insignificant male advantage in rural Eastern India in both waves. Interestingly, the small and insignificant *female* advantage apparent in the South in the early nineties becomes even smaller (in urban areas) or changes sign (in rural areas) in the second wave. The overall picture suggests very small *increases* in male advantage, both in cities, where the gap remains small and insignificant, and in rural areas, where boys' survival rates are approximately one percent higher than those for females. We will briefly comment on the bottom part of Table 2 in the next section, after we have discussed the measures of nutritional status that we will use in the rest of the paper.

3 Child Nutritional Status

The use of anthropometric indices to evaluate child nutritional status is a well-established practice (see, for example, Waterlow *et al.* (1977), WHO Working Group (1986), Gorstein *et al.* (1994)). Because weight can change in a relatively short period of time as a consequence of changes in nutritional intake and/or health status, weight-for-age and weight-for-height are better measures of short term nutritional status. However, weight-for-age cannot distinguish between small but well fed children and tall but thin ones, so that weight-for-height is usually the preferred indicator. Weight-for-height has also the advantage of not depending on age, which is frequently misreported, especially among respondents with low levels of literacy. Height-for-age is instead the preferred measure of long-term nutritional status.

To gauge the nutritional status of a child, one needs to refer to a corresponding 'normal' outcome for a child of the same age and sex. The common practice is to make use of *z-scores*, calculated as $(x_{ig} - x_g)/\sigma_g$, where x_{ig} is the weight (height) for a specific child i in group g (defined by sex and either age or height), and x_g and σ_g are respectively the mean (or median) and the standard deviation of the indicator for children within the same group in a reference population. Z-scores

⁶Such a measure of differential mortality rate was used in Rosenzweig and Schultz (1982), who were among the first to document the association between son preference and the gender gap in the economic returns to children.

are then easy to interpret if the corresponding nutritional indicator is approximately normally distributed in the reference population. If, say, a boy has a weight-for-height z-score below -1.645 then his weight is below that of 95 percent of boys in the reference population with the same height.⁷ Children are said to be *stunted* if their height-for-age z-score is below -2 , and *wasted* if their weight-for-height is below the same threshold.

Both NFHS waves also report z-scores calculated adopting the 1977 CDC growth charts for American children as a reference. These reference growth charts have been widely used as an international standard for cross-country anthropometric comparisons and their use as a reference has been recommended by the World Health Organization (Dibley *et al.* 1987a, 1987b). Such recommendations are based on evidence supporting the hypothesis that well-nourished children in different population groups follow very similar growth patterns (Martorell and Habicht (1986)). Agarval *et al.* (1991) and Bhandari *et al.* (2002), showed that these charts describe reasonably well the growth process of Indian children living in affluent families.⁸

Although changes over time of mean nutritional status can be evaluated without the use of reference growth charts, we choose to make use of z-scores because we are also interested in boy *versus* girl nutritional status. The use of z-scores facilitates such comparisons, as boys and girls have different growing patterns. Moreover, the use of z-scores is convenient because it allows one to construct a measure of nutritional status comparable across all age groups, and whose distribution can be easily tracked over time.

In the first three panels of Figure 1, we plot gender-specific nonparametric locally weighted regressions (Fan (1992)) of z-scores on age, using all valid data from NFHS-I. We emphasize that while z-scores for weight-for-age are for all India—rural and urban—the z-scores for height-for-age and weight-for-height exclude states for which height was not recorded in NFHS-I, namely Andhra Pradesh, West Bengal, Himachal Pradesh, Madhya Pradesh, and Tamil Nadu. All the patterns of the z-scores are consistent with what is commonly observed in low-income countries, and show weight and height performances which are, on average, well below those of the American children in the reference population. The curves for weight-for-age (panel A) start below zero, decline until the age of about eighteen months, and then stabilize below -2 . For both boys and girls, then, the *mean* weight performance is therefore approximately equal to that of the first percentile of the

⁷In reality, anthropometric indicators are not well described by a normal distribution. Recently revised pediatric growth charts for American children account for this, and provide an alternative method for the calculation of z-scores that still retains their interpretation in terms of quantiles of a normal distribution. For details, see Kuczmarski *et al.* (2000).

⁸However, see Klasen (1999) and Klasen and Moradi (2000) for a more skeptical view on the appropriateness of the CDC references.

reference population. Height-for-age (panel B), which represents a measure of long-term nutritional status, presents an even more striking pattern, and the regressions are still sloping downwards (and approximately equal to -3) for 4-year old children. As a consequence of low weight *and* low height, the weight-for-height z-scores (panel C) remain close to -1 , and they display positive slopes for older children, as weight-for-age stabilizes while height-for-age keeps decreasing. From these pictures, the existence of any “male advantage” in nutritional status is not apparent. In panel D of Figure 1 we pool together boys and girls and we estimate nonparametric kernel densities of z-scores for the three indexes. Consistently with the results in the previous panels, most of the probability masses remains well to the left of zero, especially for weight and height for age. Notice also that while weight and height for age appear to have similar means, the distribution of height-for-age is much more spread out.⁹ This is not surprising, as in human populations weight has a natural tendency of showing higher variability than height.

In the bottom part of Table 2 we report some preliminary evidence on child nutritional status, as described by the anthropometric indices that we have discussed in the previous paragraphs. For each indicator, we calculate gender-wave-sector specific figures for the fraction of children with z-score below -2 . All results are calculated for children up to 3 years old living in states that did not belong to Phase I of NFHS-I. The results for all India are calculated including also the smaller states and Union Territories, while the regional breakdown follows the categorization described previously. We do not discuss these results here because in the next section we will turn to the analysis of the whole distribution of these nutritional indicators. However, we would like to emphasize the fact that the number of observations used in each cell is always large (with the possible exception of the urban sector in Eastern states), so that the cumulative distribution functions will be estimated quite precisely.

In the next section we turn to a more detailed analysis of gender inequality in nutritional status, and its change between the two NFHS waves.

4 Gender Inequality in Nutritional Status and its Evolution Over Time

In what follows, rather than using z-score regressions or densities, all comparisons between genders, or over time, will be based on differences between cumulative distribution functions (*CDF*) of z-

⁹The means of weight-for-age and height-for-age are respectively equal to -1.98 and -1.87 , while the corresponding standard deviations are 1.35 and 1.72 .

scores. The CDFs’ are estimated over a grid of points by numerically integrating the densities, which we estimate nonparametrically using a biweight kernel, and choosing the bandwidth using the robust criterion proposed by Silverman (1986). To ensure comparability, we only include in the analysis children that belong to age groups and states represented in both waves. Hence, we exclude children who are more than 35 months old, or who live in states included in Phase I of NFHS-I, where height is only available in the second round of the survey.¹⁰ We are not concerned about the possible impact that selective migration may have on our results, as it is well known that the magnitude of cross-state relocation, as well as rural-to-urban migration, are very small in India. We are also not very concerned about the possibility that an important part of the any change over time in the gender gap in nutritional status is due to changes in gender-specific child mortality rates. The results in the top panel of Table 2 show that the male/female survival rates did change during the period of time we consider here, but only slightly so.

For each pair of CDFs’ (boys *versus* girls in a given wave, or NFHS-I *versus* NFHS-II for a given gender) we perform formal tests of the null hypothesis that the two distributions in the pair are equal. Given the complex survey design of both NFHS waves, a test of correct size should take into account the presence of clustering and stratification. For this reason, a standard Kolmogorov-Smirnov test for the equality of two densities estimated nonparametrically would not be appropriate. We then proceed “discretizing” the z-scores, by dividing the relevant range into intervals, and testing the null of equal distribution over the intervals by using a second-order corrected Pearson χ^2 statistic, as in Rao and Scott (1984).¹¹ The intuition behind the test is simple. If a discrete random variable has the same distribution in two different groups (in our case denoting either gender or wave), the ‘joint’ proportion of observations in the cell related to the i^{th} value and the k^{th} group should be identical—up to sampling error—to the product of the ‘marginal’ proportion of observations having the i^{th} value, and the ‘marginal’ proportion of observations in the k^{th} group. The test rejects the null when a normalized sum of the differences between joints and products of marginals is large. To avoid the presence of empty cells we divide the interval between -4 and 1—which contain most observations—into bins of length 0.5, while we create two additional bins for z-scores below -4, and above 1. We report the p -value of all the tests in Table 3, and we discuss the results in the following paragraphs.

Another interesting battery of tests would aim at detecting stochastic dominance. For instance, one could test the null that $CDF_{98/99}(z) - CDF_{92/93}(z) \leq 0$ for all z , that is, that a given distribution of z-scores in 1998-99 first-order stochastically dominates the distribution in 1992-93. This

¹⁰Results for these excluded groups are available upon request from the authors.

¹¹This test is easily implemented into STATA© using the command `svytab`.

would mean that the proportion of children whose z-score is below a certain threshold is smaller in 1998-99 than in 1992-93 for *any* value of the threshold. In principle, this could be done either using the tests proposed by Davidson and Duclos (2000), which are based on pairwise comparisons of distributions over a fixed grid of points, or with the Kolmogorov-Smirnov type tests described in Barrett and Donald (2003), which compare the distributions at *all* points. For completeness, in Table 3 we also report the results of the first-order stochastic dominance tests described in Barrett and Donald (2003), but in what follows we choose not to linger on them, for two separate reasons.¹² First, a quick look at Table 3, coupled with a visual inspection of the differences between the two distributions being compared, is sufficient to show that these tests provide ultimately the same information as the χ^2 tests for equality we have described in the previous paragraph. For instance, looking at the comparisons of distributions over time, when the χ^2 test rejects the null of equality of distribution, and the figure indicates an improvement in nutritional status, the test for stochastic dominance almost invariably rejects the null that the distribution in 1992-93 is sometimes below that in the later survey. Second, and more importantly, these tests (as well as those in Davidson and Duclos (2000)) do not have correct size when—as in our context—there is correlation among observations that belong to the same cluster.¹³

For each measure of nutritional status, Figure 2 shows wave and sector specific differences of CDFs' for all India, calculated as $CDF_{boys}(z) - CDF_{girls}(z)$, where z is a given value for the z-score. By construction, a *negative* difference implies the presence of a *male advantage*. In particular, a line everywhere below zero would indicate that the distribution for boys first-order stochastically dominates the distribution for girls, so that for every value of z-score the proportion of girls whose outcome is *below* the specified value is *higher* than the corresponding proportion for boys. Also, by the definition of a CDF, the differences approach zero when z moves towards the lower or the upper bound of the range of the relevant variable. In evaluating these differences it should also be kept in mind that most of the z-score distributions lie approximately between -4 and zero, with medians close to -2 for weight and height for age, and around -1 for weight-for-height. For this reason, the emphasis of our analysis will be on the behavior of the lines evaluated at negative z-scores.

Several interesting patterns emerge from the graphs in Figure 2. In 1992-93 (continuous lines)

¹²In comparing two distributions $F(z)$ and $G(z)$, a statistic such as those described by Barrett and Donald (2003) tests the null $G(z) \leq F(z)$ for all z , versus the alternative that $G(z) > F(z)$ for *some* z . The statistic is calculated at $\sqrt{NM/(N+M)} \sup_z (\hat{G}(z) - \hat{F}(z))$, where N and M refer to the size of the two samples used to produce the two estimates $\hat{G}(z)$ and $\hat{F}(z)$.

¹³According to our calculations, intracluster correlation for z-scores is always between .1 and .18, for all measures of nutritional status, and for either sector in both NFHS rounds. Monte Carlo simulations performed by the authors show that size distortion can be large in presence of intracluster correlation.

there is no clear evidence of generalized female disadvantage in nutritional status, as evaluated relative to US growth charts. In fact, in both sectors and for almost all values of z , *growth performances appear to be relatively better for girls*. For example, in rural areas the proportion of girls whose weight-for-age is below -2 is about 2 percentage points lower than for boys, even if at conventional levels the χ^2 test does not reject the null that the two gender specific distributions are equal. Looking at weight-for-height, the gap is almost twice as large, and the test of equality of distribution strongly rejects the null. In urban areas differences are generally very small, especially for height-for-age, and not statistically significant. However, the curves calculated from NFHS-II (dashed lines), show a clear and striking change, especially in rural areas, where in 1998-99 all curves lie virtually everywhere *below* the corresponding curves in the previous NFHS wave, indicating a clear movement towards male relative advantage in nutritional status. In NFHS-II, the proportion of girls whose weight-for-age z-score is below -2 becomes about 3 percentage points *higher* than for boys, and the χ^2 test now rejects the null that the distribution for weight-for-age is the same across genders. The difference in the proportion of stunted children preserves the same magnitude as in NFHS-I (the test still does not reject the null of equality of distributions) but the sign is reversed, and the female advantage in weight-for-height almost completely disappears, and becomes a small male advantage over part of the range. The picture for urban areas is quite different, and shows relatively gender-neutral results in both survey rounds, although it is still true that the curves for 1998-99 lie almost everywhere below the corresponding lines for 1992-93. However, in no case is the null rejected using a 5% significance level.

To analyze whether the change in the gender difference in distributions is not only large in magnitude but also statistically significant, in Figure 3 we plot sector-specific “differences-in-differences” of CDFs’ for all India, together with 95% confidence bands. For brevity, we only consider weight and height for age.¹⁴ The difference-in-differences is calculated as

$$[CDF_{boys}^{II}(z) - CDF_{girls}^{II}(z)] - [CDF_{boys}^I(z) - CDF_{girls}^I(z)],$$

where the superscripts denotes the NFHS wave. Because “relative boy advantage” translates into *negative* values of each difference, an *increase* in boy advantage will be represented by a *negative* difference-in-differences. We construct 95% confidence bands using bootstrap, with 250 replications. In each replication, and independently for rural and urban areas, we first resample *clusters* separately from each NFHS round. We then re-estimate all the difference-in-differences at each replication and calculate the value of the lower and upper bands for each point on a grid as the 2.5 and 97.5 percentiles from the bootstrap distribution. Because resampling with clusters in-

¹⁴The results for weight-for-height are available upon request.

cludes all observations for a selected cluster, this procedure takes into account both intracluster and intra-household correlation, so that the confidence intervals should have correct coverage rates. The confidence bands in Figure 3 show that in rural areas, and especially for weight-for-age, the increase in boy advantage is large, and for most of the relevant range the upper band lies below zero, indicating that over such range the null hypothesis of no change in the gender gap would be rejected at a 5% significance level. In urban areas the difference-in-differences over the relevant range is also negative, but in this case the bands include zero throughout the whole range.

Clearly, the results discussed so far are silent about the *absolute* change in nutritional status during the period. In Figure 4, we turn to analyze explicitly such changes. Each graph plots the two sex-specific changes over time of the corresponding CDFs'. Because each line is calculated as $CDF_{98-99}(z) - CDF_{92-93}(z)$, *improvements* in relative nutritional status are represented by *negative* values. Stark gender differences emerge, especially in rural areas. Among children living in the countryside, large improvements in weight-for-age performances are only apparent among boys: while the proportion of boys whose z-score is below -2 decreased by about 7 percentage points between the two waves, we only observe a less than 2 percent drop in the corresponding figure for girls. Indeed, the change in CDF for girls is everywhere close to zero, and the p-value of the χ^2 test is 0.046, so that the null of no change over time is barely rejected using a 5% significance level. The p-value of the test for boys is instead approximately zero. Looking at height-for-age, two results stand out. First, girl performances appear to have slightly *worsened*, during this period, a result which is difficult to reconcile with the small but still positive change in weight-for-age, although one cannot reject the null of no change over time using standard significance levels. Second, height performances for boys have instead improved, but not as much as for weight-for-age, so that the p-value of the χ^2 test is approximately 0.06. The fact that changes in height appear to be much less favorable than changes in weight may be a result of the fact that the former is much less responsive than the latter to changes in health and nutrition, and it is also more likely to be heavily influenced by genetic factors. Weight-for-height shows improvements for both genders, once again larger for boys than for girls, but it should be noted that the improvement for girls is just a result of small improvements in weight-for-age and small *worsening* in height-for-age. These results also highlight the importance of looking at different indicators when evaluating changes in the nutritional status of a population, as important elements may be missed otherwise.

The bottom three panels in Figure 4 show that changes appear to be more similar between genders in urban areas, a result confirmed by the difference-in-differences discussed before. Improvements in boy weight-for-age are large and similar to those observed in rural areas, but girls appear to be doing much better as well: for example, the proportion of girls with z-score below -2

decreased by 4 percentage points between 1992-93 and 1998-99, while the proportion decreases by about 5 percent for boys. However, the gender gap in the change widens for z-scores between -2 and -1 . For both genders, the χ^2 test rejects the null of no change over time in the distributions, even using a 1% level. Changes in height-for-age are more unequal: while $CDF(z)$ decreases for girls by about 2 points for z in the interval between -4 and -2 , the drop for boys is twice as large, or larger. The different changes for weight and height compound to generate large improvements in weight-for-height, improvements who appear to be remarkably similar between genders.

Overall, our analysis shows important changes in the distribution of anthropometric indicators during the relatively short six-year period between the two NFHS surveys. But while boy weight and height performances are significantly better in the later period, improvements are generally much less clear-cut for girls, and height-for-age even shows some worsening for girls living in rural areas. In the next section we turn our attention to analyzing whether there are important—and changing—geographical patterns in these anthropometric indicators.

4.1 Region and sector specific results

It is well known that there are important differences in the extent of gender inequality in different regions of India. For this reason, it is useful to study the possible existence of geographical patterns in boy *versus* girl nutritional status. Ideally, it would be interesting to conduct separate analysis for each state, and for rural versus urban areas. However, in order to preserve a relatively large number of observations in each area, we work with three broadly defined regions—North, East, and South—and we consider the rural and the urban sector separately in each region. The regional breakdown follows the geo-cultural classification proposed by Sopher (1980) and accepted, amongst others, by Bourne and Walker (1991), Dasgupta (1993) and Dyson and Moore (1983). As in the previous section, for comparability reasons we only include children up to 3 years old, and we exclude from the analysis the states for which height is missing in NFHS-I. The remaining states are then grouped as follows: North combines Gujarat, Haryana, Jammu, Punjab, Rajasthan, Uttar Pradesh, and New Delhi; East is composed of Assam, Bihar, and Orissa, while Kerala, Karnataka and Maharashtra represent the South. For conciseness, we only consider weight and height for age. The previous section has shown that information on weight-for-height can be easily deduced from the results obtained for these two indicators.¹⁵

In the top four panels of Figures 5a, 5b and 5c, we plot wave and sector specific gender gaps between CDFs' of weight and height for age z-scores for North, East, and South India respectively.

¹⁵Results for weight-for-height are available upon requests.

In the bottom four panels we display the changes over time of gender and sector specific CDFs' for the same indicators. Figures 6a to 6c include the differences-in-differences estimates, as well as the corresponding 95% confidence bands. The p-values of the χ^2 tests for equality of each pair of distributions (including those for weight-for-height) are again reported in Table 3. Several striking differences are apparent, both across different regions and between rural and urban areas within the same region.

The results for North India are relatively similar to those for the whole country discussed in the previous section. This is probably not surprising, as the North accounts for approximately sixty percent of the total observations in the sample. Looking at the top 4 graphs in Figure 5a, we note that in both urban and rural areas, girl nutritional status in NFHS-I appears to be overall relatively *better* than for boys (although the null of equal distribution of z-scores is not rejected for both weight and height at the 5% significance level). However, the curves for 1998-99 lie almost everywhere below zero, indicating an overturn of the relation which is especially clear in rural areas, and for height in urban areas. For example, looking at weight-for-age in rural areas, the difference in boy *versus* girl proportion with z-score below -2 goes from one in 1992-93 to -5 in the following wave. This change in the gap is statistically significantly different from zero, as shows by figure 6a. Using data from the rural sample of the second round of the NFHS, the χ^2 test rejects the null of equal distribution for both weight and height. The bottom four graphs in Figure 5a show that this is the result of a virtually unchanged distribution of weight-for-age z-score for girls, accompanied by a fairly large improvement for boys (e.g. $CDF_{boys}(-2)$ decreases by approximately 6 percentage points). Weight performances in urban areas improve considerably both for boys and—to an only marginally smaller extent—for girls. Changes in the distribution of height-for-age present a different picture. In urban areas we still observe improvements, but much less pronounced than for weight-for-age, especially for girls, for whom the difference in CDFs' is everywhere very close to zero. In the countryside our results indicate a *worsening* of height performances for both genders, marginally so for boys, but surprisingly large for girls: the proportion of girls with z-score below -2 appear to have *increased* by about five percent among girls, and the p-value of the χ^2 test is 0.0004. The difference-in-differences in bottom-right panel of Figure 6a confirm a generalized movement towards boy advantage in height-for-age z-scores in rural areas, even if zero is included in the 95% confidence bands over most of the range. The pattern in the urban sector is very similar.

In the Eastern region as well (Figure 5b), we observe generalized improvements in boy nutritional status relative to girls. This phenomenon is particularly stark in rural areas, as confirmed by the differences-in-differences, in Figure 6b. In 1992-93, the distribution of girl weight and height for age appears to stochastically dominate the corresponding distribution for boys (and the χ^2

test always rejects the null of equal distributions), but the relation between CDFs' is overturned throughout the whole relevant range in the next wave, so that in the later survey one never rejects the null of equal distributions across genders at standard significance levels. The same happens to weight-for-age in urban areas, where, for example, $CDF_{boys}(-1) - CDF_{girls}(-1)$ moves approximately from 0.04 in 1992-93 to -0.06 in 1998-99. The top-right panel of Figure 6b shows that the difference is statistically significant, even if zero is included between the (fairly wide) confidence bands throughout most of the range, partly because of the relatively small sample size (see Table 2). Only for height-for-age in urban areas, and for z-scores below -1 we do not observe a large increase in "boy advantage", but this is also the only case where a clear male advantage was already apparent in the earlier period. Looking at the four bottom graphs of Figure 5b, we see that weight and height performances in rural areas improved considerably over the period for boys, but remained almost unchanged for girls. In urban areas, the data suggest instead large reductions in the proportion of boys *and* girls with height and weight z-scores below very low thresholds, but an increase in the gender gap in weight-for-age is clear for z-scores between -2 and 0 .

The picture for the South is completely different.¹⁶ The graphs in the top part of Figure 5c show that in urban areas there is *no* evidence of a movement towards male advantage in nutritional status. Indeed, over certain ranges of weight and height z-scores the differences in CDFs' move upwards. In both waves and for both indicators the data suggest the existence of a girl advantage which does not appear to waver over time, even if the p-value of the χ^2 test for equality of distributions is above 0.10 in all cases. In rural areas, the pattern of change in gender inequality is somewhat more consistent with what we observe in other regions, as for negative values of z-scores the lines move generally downwards, although the vertical distance between the wave-specific lines is never larger than—and is mostly much lower than—4 percentage points. The differences-in-differences confirm a very stable relation between boy and girl nutritional status, and nowhere in the relevant range is zero outside the 95% pointwise confidence bands. The bottom four graphs in Figure 5c show that the relative stability of gender inequality in height and weight is the result of generalized improvements over time. Notice, however, that the changes over time are relatively small, and using standard significance levels the χ^2 test only rejects the null of no change over time in the case of height-for-age in rural areas.

Overall, we observe important movements in the distribution of weight and height for age z-scores during the short period of time between the two waves of the NFHS, but we also find that in rural areas of Northern and Eastern states boys appear to have benefitted much more than girls from a period of rapid economic growth. Only in Southern regions, and in urban areas elsewhere,

¹⁶Remember that here the South only includes Kerala, Karnataka and Maharashtra.

do we find clear improvements in the nutritional status of children (up to 3 years old) for both boys and girls. It is somehow disturbing that areas where son preference has historically been found to be stronger—and in a period of rapid growth—appear to be moving towards a situation of *more* pronounced gender inequality in child nutritional status.

5 Nutritional Status and Socioeconomic and Demographic Factors

In this section we evaluate how much of the changes in the distribution of z-scores described above can be explained by changes over time in the distribution of a list of socioeconomic and demographic factors that are likely to be important predictors of child nutritional status. For this purpose, we use an approach borrowed from DiNardo, Fortin, and Lemieux (1996), which is a semiparametric analogue to the more familiar Oaxaca decomposition for linear regression models (Oaxaca (1973)). We ‘decompose’ the change in the cumulative distribution function $F(z)$ of an anthropometric index z using the following thought experiment: what would the marginal distribution of z look like (in 1998-99) if $F(z | \mathbf{x})$, was the same as in 1992-93 but $F(z)$ was as in 1998-99? We are particularly interested in evaluating the contribution to changes in nutritional status of variables such as income and parental education.

The construction of *counterfactual* distributions such as the one described above proceeds using a semi-parametric estimator that we describe below, but two caveats should be immediately highlighted. First, this approach does *not* pretend to uncover a *causal* relationship between the included covariates and nutritional status, as many of the variables included in \mathbf{x} would be clearly endogenous in a regression context. In fact, child nutrition and health care are surely jointly determined with many other decisions at the household level, such as fertility and labor supply. Second, even if the counterfactual distribution for the second NFHS wave were identical to the actual one, one cannot take this as evidence that the conditional density $f(z | \mathbf{x})$ has not changed over time. In fact, for a given $f(\mathbf{x})$, it is not difficult to see that the relationship between $f(z | \mathbf{x})$ and $f(z)$ is not invertible.¹⁷

In the previous sections, we have described how most anthropometric indices show sharp improvements during the nineties. Because this period was characterized by rapid economic growth, it is natural to expect that increases in household resources explain a sizeable fraction of such improvements. Unfortunately, the NFHS does not record information on income or consumption, but

¹⁷Indeed, there are reason to think that the conditional density will change over time, as the “technology” that links nutrition, health care, and epidemiological environment is unlikely to remain stable, and the relatively parsimonious list of variables that we include in \mathbf{x} is likely to omit some important factors.

the household questionnaire does include questions on several wealth indicators, including dwelling characteristics and ownership of a long list of items. This allows us to construct an *asset index* that we use proxy for the economic status of the household. We construct the index using principal component analysis, a practice that is now commonly adopted in the presence of data limitations such as the ones we face here. For example, Filmer and Pritchett (2001), using NFHS data, show that an asset index constructed using principal component analysis performs well as a proxy for long-run economic status, and argue that such an index might be a better measure of household wealth than current consumption or income.¹⁸ The appendix contains a detailed description of the construction of the asset index.

Besides the asset index, we include in the vector of covariates \mathbf{x} mother and father’s education (in the form of dummies for no education or at least high school diploma), child age, household size, number of older siblings, and we finally include individually some of the variables used to construct the asset index, as they may have a direct impact on child health status beyond their contribution to the construction of the asset index. Such variables are dummies indicating the main source of drinking water and the availability of electricity or sanitation facilities in the house.

We turn now to the description of the estimation of the counterfactual distributions. Formally, let $f(\bar{z} | t)$ be the true density of the anthropometric index z evaluated at \bar{z} , in wave t , where $t = I, II$. The density can be rewritten as¹⁹

$$f(\bar{z} | t) = \int f(\bar{z} | \mathbf{x}, t) f(\mathbf{x} | t) d\mathbf{x}$$

where $f(\mathbf{x} | t)$ is the density of the covariates \mathbf{x} in wave t . For notational convenience, let us write $f(\bar{z} | t) \equiv f(\bar{z} | t_{\mathbf{x}} = t, t_{z|\mathbf{x}} = t)$, where $t_{z|\mathbf{x}}$ indicates the wave that identifies the conditional distribution of z given \mathbf{x} , and $t_{\mathbf{x}}$ indicates the wave that identifies the marginal distribution of \mathbf{x} . Clearly, the two waves coincide in the *actual* density. We are interested in studying if the changes in the distribution of z-scores can be completely explained by changes in the distribution of the covariates, *assuming that the conditional distribution of $z | \mathbf{x}$ does not change*. This amounts to estimating the counterfactual density $f(\bar{z} | t_{\mathbf{x}} = II, t_{z|\mathbf{x}} = I)$. A straightforward way to estimate this object follows after noting that it can be usefully rewritten as follows (see Appendix B for a proof):

$$f(\bar{z} | t_{\mathbf{x}} = II, t_{z|\mathbf{x}} = I) = f(\bar{z} | t = I) E[R(\mathbf{x}) | \bar{z}, t = I] \tag{1}$$

¹⁸But see Mukherjee (2004) for a critique of this methodology, and for an alternative technique to use asset ownership indicators to construct a proxy for household wealth. Mukherjee (2004) also presents extensive references to other recent papers that use principal components analysis for analogous purposes.

¹⁹We describe the estimation for the case where the covariates \mathbf{x} are continuous, but with a change of notation the argument can be straightforwardly adapted to the case where at least some of the covariates are discrete.

where

$$R(\mathbf{x}) = \frac{P(t_{\mathbf{x}} = II | \mathbf{x}) P(t_{\mathbf{x}} = I)}{P(t_{\mathbf{x}} = I | \mathbf{x}) P(t_{\mathbf{x}} = II)}.$$

The function $R(\mathbf{x})$ is a *reweighting function* that maps the conditional density from wave I into the counterfactual density $f(\bar{z} | t_{\mathbf{x}} = II, t_{z|\mathbf{x}} = I)$, by increasing (decreasing) the contribution to this counterfactual marginal density of the conditional density $f(\bar{z} | \mathbf{x}, t_{z|\mathbf{x}} = I)$ for values of \mathbf{x} that are relatively common (rare) in wave II . To interpret the different components of the reweighting function it is useful to consider a “super-sample” that encompasses data from both NFHS waves. Then the unconditional probability $P(t_{\mathbf{x}} = I)$ can be simply estimated as the (weighted) fraction of observations that belongs to the first wave, while the conditional probability $P(t_{\mathbf{x}} = II | \mathbf{x})$ can be interpreted as the probability that an observation with covariates equal to \mathbf{x} belongs to the second NFHS wave. Because $t_{\mathbf{x}}$ is either I or II , such probability corresponds to the regression over \mathbf{x} of a binary dependent variable equal to one if the child has been sampled from the second wave. Here we estimate the regression using a flexible logit model where all continuous variables in \mathbf{x} are entered as cubic B-splines with five equally spaced knots.

The counterfactual density can then be estimated using a simple two-step procedure: first an estimate of $\hat{R}(\mathbf{x})$ is obtained and then the counterfactual density is estimated using a modified nonparametric kernel density estimator as in the following expression

$$\hat{f}(\bar{z} | t_{\mathbf{x}} = II, t_{z|\mathbf{x}} = I) = \sum_{i \in I} w_i \hat{R}(\mathbf{x}_i) \frac{1}{h} K\left(\frac{\bar{z} - z_i}{h}\right),$$

where w_i is the sampling weight for the i^{th} observation (normalized so that $\sum_{i \in I} w_i = 1$), $K(\cdot)$ is a standard kernel, h is the bandwidth and $i \in I$ indicates that the summation is taken only over observations that belong to the first wave. Once the densities have been estimated, the cumulative distribution functions can be calculated as usual by numerical integration.²⁰

For the sake of brevity, here we only report the results for all Indian states (excluding as usual states in Phase I of NFHS-I) for height and weight-for-age. We plot the resulting predicted changes by sector in Figure 7 together with the actual ones. In rural areas, the predicted change in the distribution of weight-for-age z-scores for boys is considerably smaller than the actual one (top left panel). For instance, while the predicted reduction in the proportion of boys with z-score below -2 is 3 percent, the actual reduction is more than twice as large. Interestingly, the predicted change for girls is very similar to (but slightly smaller than) the predicted change for boys. However, the actual change in this case is *smaller*, so that weight performances for rural

²⁰In principle, one can estimate directly the counterfactual CDFs’ using a procedure analogous to that just described (see Tarozzi (2004) for details). We choose to estimate the densities first because the resulting graphs are much smoother.

girls is not much different between the two rounds of the NFHS. In urban areas, the few variables included in \mathbf{x} explain a large fraction of the change for both boys and, to a lesser extent, girls. Note that, unlike in rural areas, predicted improvements in girl nutrition are overall smaller than the actual ones, and that the predicted improvements for boys are once again larger than for girls. Almost all the predicted change in the distribution of z-scores is related to improvements over time in educational achievement and asset ownership. Indeed, the estimation of counterfactual distributions where we only include demographic variables in the vector \mathbf{x} leads to the prediction of almost no change in nutritional status between the two rounds (and so we do not include them here). This latter observation is also interesting because it suggests that a change in the distribution of demographic variables (such as age or family composition) is very unlikely to be driving the changes in distributions that we detect in the data. As an aside, we also observe that the fact that predicted improvements are larger for boys than for girls suggests the existence of a stronger association for boys than for girls between education/asset ownership and nutritional status.

It is probably not surprising that part of the change in the distributions of z-scores is not explained by the few covariates we have included in our analysis. In particular, the asset index is likely to be an imprecise measure of wealth, and we are not taking into account other factors that are likely to have benefited child nutritional status, such as improvements in the overall epidemiological environment which we would expect during a period of rapid economic growth such as the nineties. Rather, the striking result is that in rural areas girl weight appears to have improved *less* than predicted by our thought experiment. The same result is found, only more pronounced, when we look at the predicted changes in the distribution of height (bottom four graphs in Figure 7). In rural areas, the counterfactual distributions predict improvements for both genders. However, as we have already observed, for girls the data show a *deterioration* of height performances. In urban areas, the predicted and actual changes in girl height are instead quite close. The patterns of actual and predicted changes are also very close when we look at the distributions for boys, in both sectors. However, over most of the relevant range, the predicted change in the rural distribution of boy height z-scores now *overstate* the actual improvement, while the opposite was true for weight. Notice also that the predicted improvements for boys are once more larger than those we observe for girls, and they are overall slightly smaller than the predicted changes in weight.

6 Conclusions

In this paper we have described some new interesting stylized facts on the changes in boy versus girl nutritional status in India in the years between 1992-93 and 1998-99, a period during which

the rate of growth of real per capita GDP remained constantly above 4 percent per year. Overall, we find that the nutritional status of young children (age 0-3)—as measured by weight and height performances relative to those of a reference population—improved, but we also find that gender differences in nutritional status increased as well, with nutritional status improving substantially more for boys than for girls. Consistent with a large literature that documents the existence of a steep North-South gradient in gender inequality in India, we find that improvements in nutritional status appear to be gender neutral only in the South. In particular, in the North and East we find that girl nutritional status has not improved in the countryside, and height performances even appear to have worsened. We also estimate predicted changes in nutritional status based on changes in the distribution of household wealth (proxied by asset ownership) and a few other observed household characteristics. Actual changes appear to be relatively close to predicted ones, especially in urban areas, although actual changes are somewhat higher than predicted ones for boys in rural areas. We estimate that the predicted changes are generally larger for boys than for girls.

In the introduction, we have discussed reasons why one should not necessarily expect a reduction in son preference as a consequence of economic development, even when son preference is at least partly related to the existence of economic constraints. Our results suggest that in certain areas of India it may be indeed the case that improvements in parental wealth and educational level have benefited boys much more than girls. The relative large samples available in the surveys we have used, the result of formal tests of statistical significance, as well as the fact that we do not observe important gender disparities in the changes in the South (where son preference is less pronounced), lead us to think that the results we document are not simply due to sampling error. It would be useful to corroborate our result with different data sources, and it will be very interesting to study if analogous trends are observed in the third round of the National Family and Health Survey, which at the time of writing is being conducted in the field.

Clearly, two questions arise naturally from our results: why do we observe, at times, such different trends in height and weight performances, and—above all—why do we observe these trends in gender differences? These are important questions, especially if our results are confirmed by other independent data sources, and we hope to analyze such issue in future research.

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Appendix A - The construction of the asset index

Principal component analysis is a statistical procedure whose objective is to explain the variance of a set of k variables (which here represent asset ownership) in terms of k orthogonal factors (or principal components) whose meaning should be suggested by the context. Each principal component is calculated as a particular weighted average of the k —standardized—variables (details on principal component analysis can be found, for example, in Lindeman, Merenda, and Gold (1980)). The largest share of the total variance is usually explained by the first principal component, which is then interpreted as a “logical common denominator” of the k variables. In the present context such variables measure asset ownership, so that the first principal component is naturally interpreted as a measure of wealth.

The variables that record asset ownership in the two waves of the NFHS are not identical, so that to ensure comparability we only make use of those common to the two datasets. This leaves us with a list of 22 variables which include the following: dummies for ownership of television, radio, clock or watch, fan, sewing machine, refrigerator, bicycle, motorcycle, car, livestock; indicators for whether the household uses a separate room used as kitchen, the house is of high quality or low quality materials, the main cooking fuel is of “high quality” (kerosene, electricity, or L.P.G.), the main source of drinking water is piped into the residence, or it is surface water instead, the house has flush toilet facilities, or no toilet, the house has electricity; finally, we include the number of rooms in the house, the amount of land cultivated, and of land irrigated and owned by the household. When one or more of the asset ownership variables are missing, we impute a value equal to the district median for the variable. This imputation procedure is only necessary for less than one percent of the observations, and in almost all cases only one imputation is necessary, so that the results are virtually unchanged if we exclude all incomplete observations. Because the association between wealth and the ownership of different assets is likely to be different in rural and urban areas, we calculate separate sets of weights for the first principal component in the two sectors. The estimated weights are reported in Table A1. Note that because the first principal component is calculated as a weighted average of the standardized original variables, the ratio between the weight and the standard deviation represents the change in the index associated with a unit change in the original variable.

Appendix B - Proof of equation (1)

The proof follows from a straightforward application of the properties of probabilities.

$$\begin{aligned}
 f(\bar{z} | t_{\mathbf{x}} = II, t_{z|\mathbf{x}} = I) &= \int f(\bar{z} | \mathbf{x}, t_{z|\mathbf{x}} = I) f(\mathbf{x} | t_{\mathbf{x}} = II) d\mathbf{x} \\
 &= \int f(\bar{z} | \mathbf{x}, t_{z|\mathbf{x}} = I) \frac{f(\mathbf{x} | t_{\mathbf{x}} = I)}{f(\mathbf{x} | t_{\mathbf{x}} = I)} f(\mathbf{x} | t_{\mathbf{x}} = II) d\mathbf{x} \\
 &= \int \frac{f(\mathbf{x} | t_{\mathbf{x}} = II)}{f(\mathbf{x} | t_{\mathbf{x}} = I)} f(\bar{z}, \mathbf{x} | t = I) d\mathbf{x} \\
 &= f(\bar{z} | t = I) \int \frac{f(\mathbf{x} | t_{\mathbf{x}} = II)}{f(\mathbf{x} | t_{\mathbf{x}} = I)} f(\mathbf{x} | \bar{z}, t = I) d\mathbf{x} \\
 &= f(\bar{z} | t = I) E[R(\mathbf{x}) | \bar{z}, t = I]
 \end{aligned}$$

where the last step follows noting that

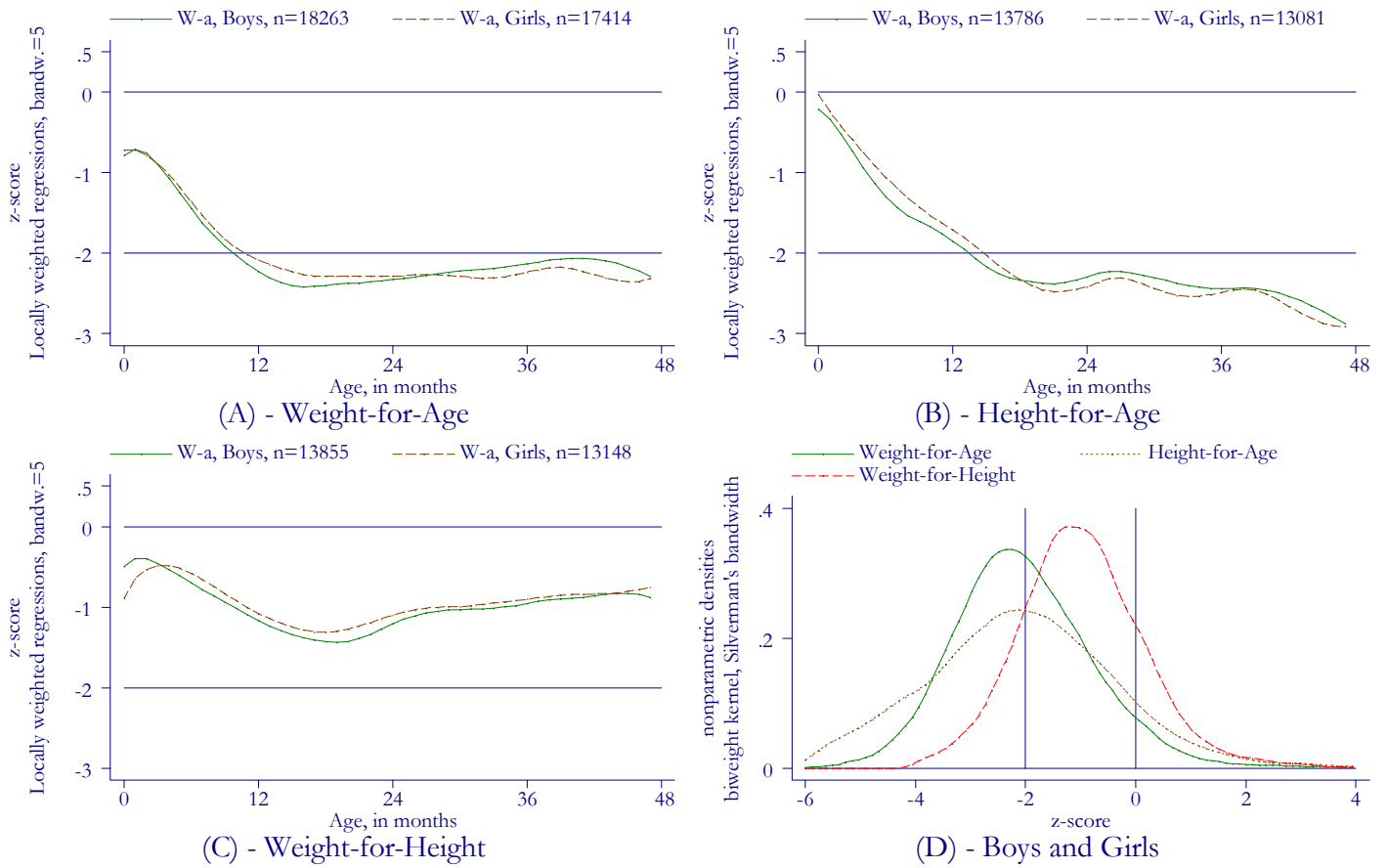
$$\begin{aligned}
 \frac{f(\mathbf{x} | t_{\mathbf{x}} = II)}{f(\mathbf{x} | t_{\mathbf{x}} = I)} &= \frac{f(\mathbf{x}, t_{\mathbf{x}} = II)}{P(t_{\mathbf{x}} = II)} \frac{P(t_{\mathbf{x}} = I)}{f(\mathbf{x} | t_{\mathbf{x}} = I)} \\
 &= \frac{P(t_{\mathbf{x}} = II | \mathbf{x}) f(\mathbf{x})}{P(t_{\mathbf{x}} = I | \mathbf{x}) f(\mathbf{x})} \frac{P(t_{\mathbf{x}} = I)}{P(t_{\mathbf{x}} = II)} \\
 &= R(\mathbf{x}) \quad \blacksquare
 \end{aligned}$$

Table A1 - Construction of the Asset Index

	Rural Sector		Urban Sector	
	Weight	Weight/s.d.	Weight	Weight/s.d.
Owns a telephone	0.299	0.6215	0.312	0.8171
Owns a radio	0.205	0.4164	0.229	0.4801
# rooms in house	0.200	0.1017	0.229	0.1201
Separate room used as kitchen	0.214	0.4526	0.190	0.3798
Land cultivated (acres)	0.017	0.0002	0.067	0.0009
Land irrigated (acres)	0.016	0.0003	0.068	0.0014
Owns livestock	-0.070	-0.1928	-0.006	-0.0130
Owns clock/watch	0.245	0.7136	0.267	0.5366
Owns a fan	0.288	0.6664	0.317	0.7000
Owns sewing machine	0.224	0.4548	0.231	0.6214
Main cooking fuel Kerosene/electricity/LPG	0.256	0.5436	0.242	0.9048
Drinking water piped into residence	0.240	0.4804	0.183	0.6009
Main source of drinking water river/lake/etc.	-0.068	-0.5235	-0.038	-0.1411
Flush toilet in house	0.271	0.5417	0.241	0.8761
Has electricity	0.246	0.7895	0.268	0.5357
No toilet facility in house	-0.242	-0.6147	-0.230	-0.5497
High quality housing	0.261	0.5352	0.258	0.6908
Low quality housing	-0.222	-0.6748	-0.228	-0.4579
Owns a refrigerator	0.263	0.5790	0.234	1.1919
Owns a bicycle	0.110	0.2198	0.128	0.2633
Owns a motorcycle	0.224	0.5318	0.223	0.9690
Owns a car	0.126	0.5701	0.106	1.3118

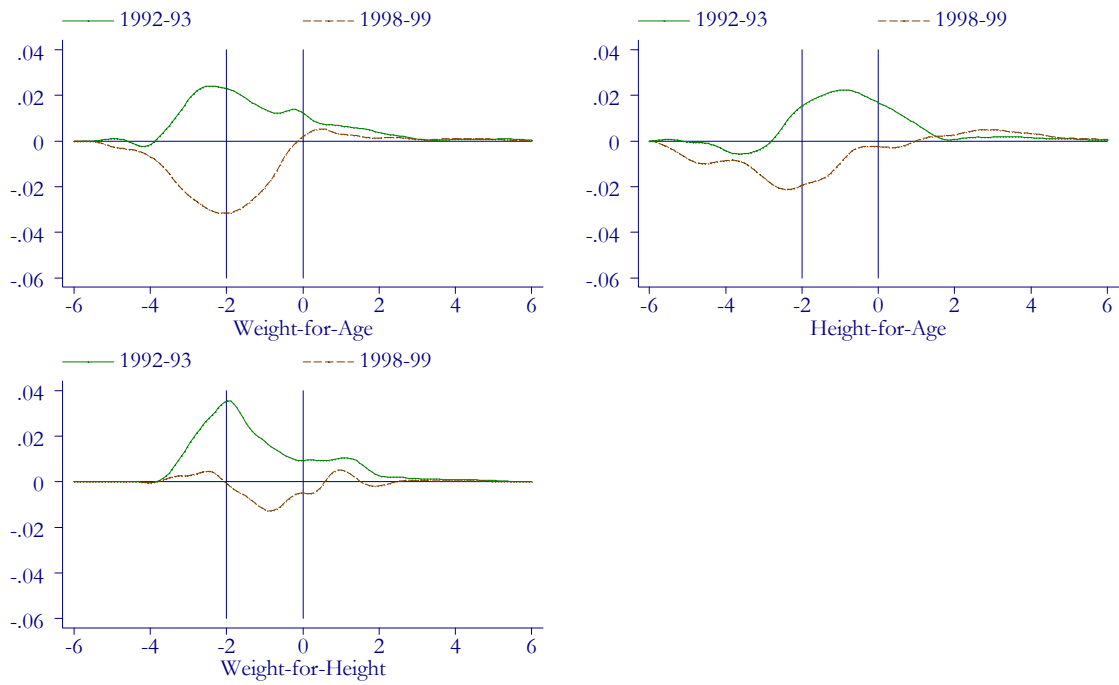
Source: authors' calculations from NFHS-I and NFHS-II.

The weights are unweighted estimates calculated pooling all available sector-specific observations from both waves. The first principal component is interpreted as a proxy for wealth, and it is calculated as a weighted average of the standardized original variables, so that the ratio between the weight and the standard deviation represents the change in the index associated with a unit change in the original variable.

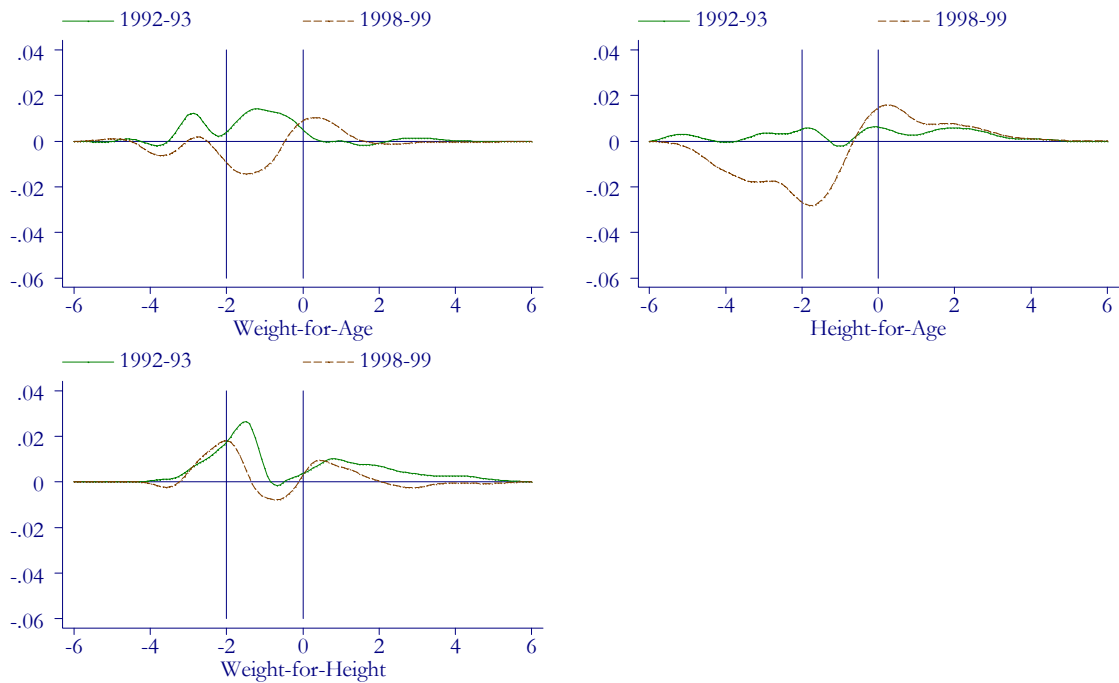


NFHS-I, 1992-93

Figure 1 – Source: authors’ calculations from NFHS-I, 1992-93. Z-scores for weight-for-age are for Indian states, rural and urban areas, while z-scores for height-for-age and weight-for-height exclude states for which height was not recorded in 1992-93, namely Andhra Pradesh, West Bengal, Himachal Pradesh, Madhya Pradesh, and Tamil Nadu.



age 0-35 months, exclude AP/HP/MP/TN/WB
Differences CDF(boys)-CDF(girls), RURAL



age 0-35 months, exclude AP/HP/MP/TN/WB
Differences CDF(boys)-CDF(girls), URBAN

Figure 2 – Boys vs. girls nutritional status. Source: authors' calculations from NFHS-I and NFHS-II. All z-scores are calculated excluding states for which height was not recorded in 1992-93 (Andhra Pradesh, West Bengal, Himachal Pradesh, Madhya Pradesh, and Tamil Nadu).

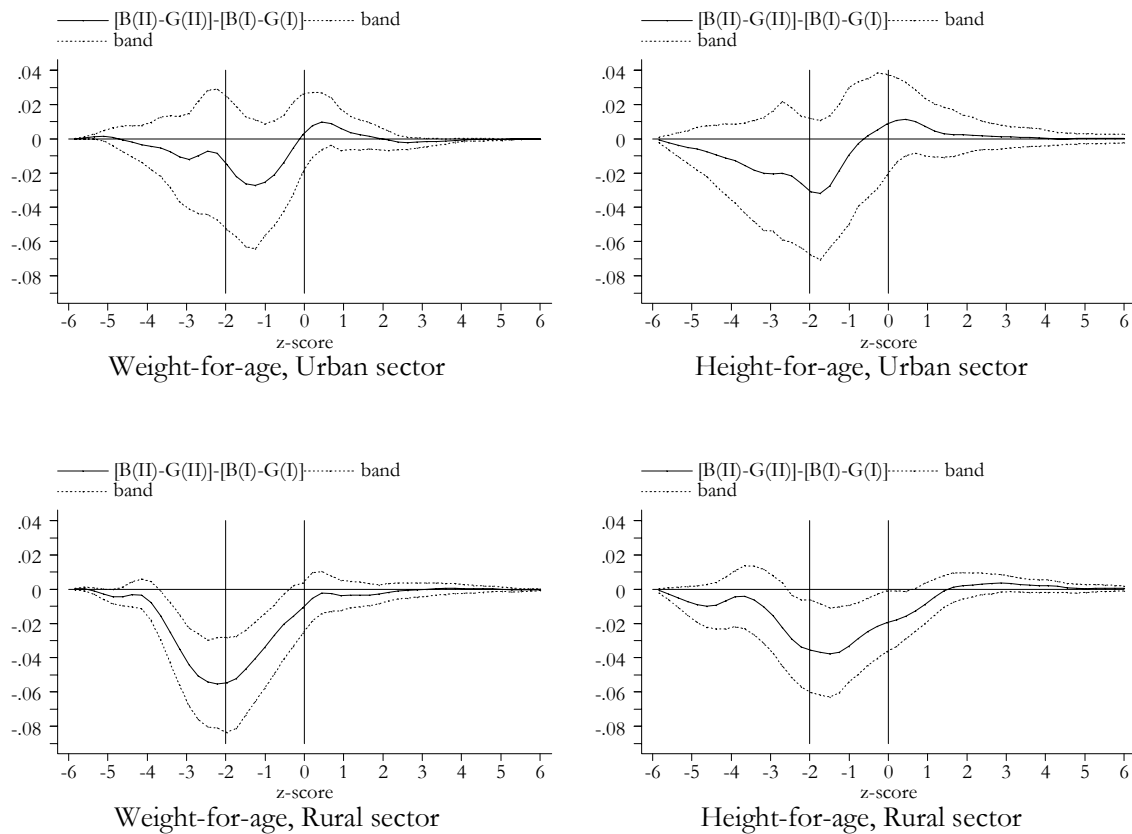
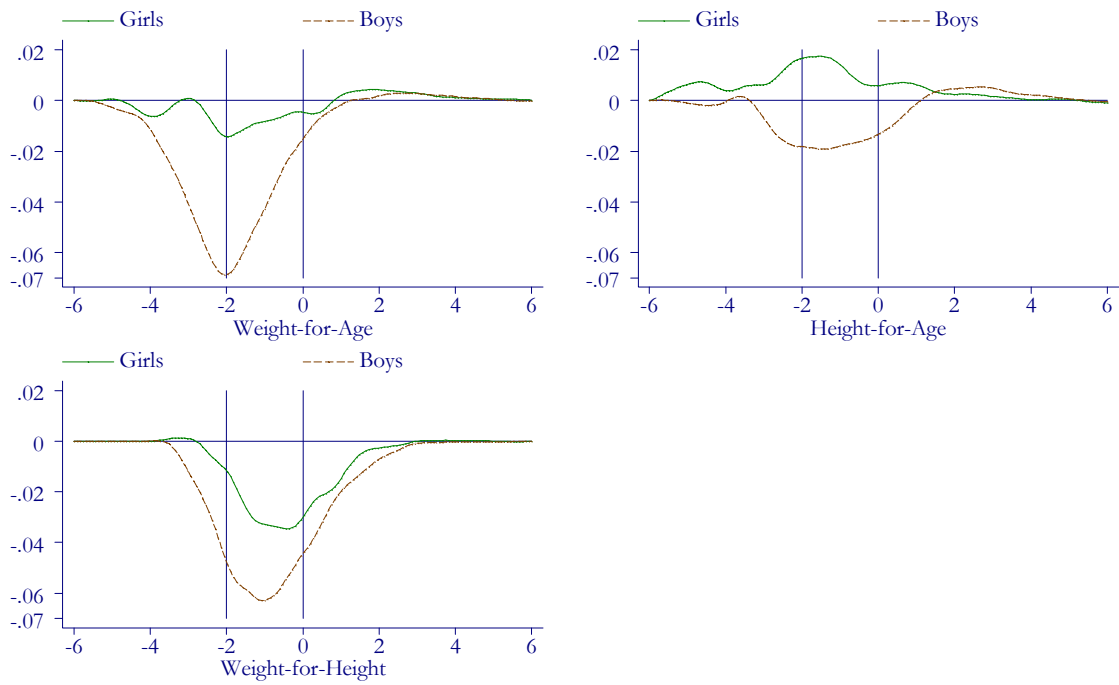
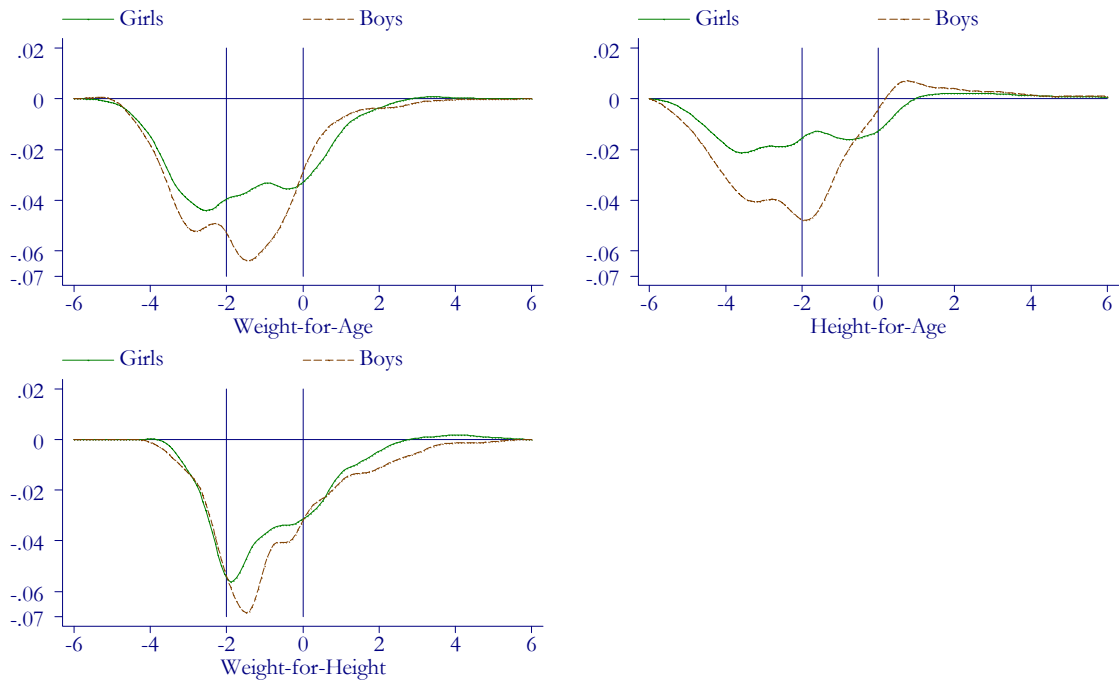


Figure 3 – All India. Source: authors' calculations from NFHS-I and NFHS-II. Children below 3 years old. States for which height was not recorded in 1992-93 are excluded. Each continuous line represents the change over time of the pointwise gender difference in distributions. By construction, negative differences-in-differences (dd) indicate an in(de)crease in boy ad(disad)vantage. The dotted lines represent 95% confidence bands, calculated, at each point over the grid, as the 2.5 and 97.5 percentile of the distribution of estimated dd over 250 bootstrap replications, each done by resampling clusters separately from each survey.

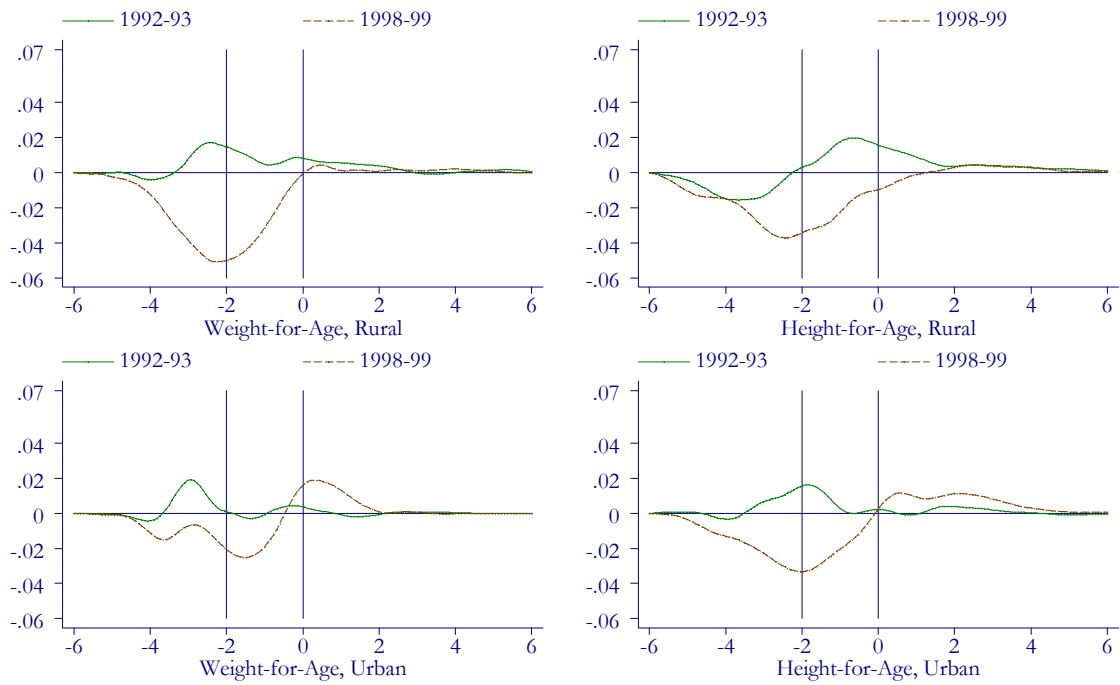


age 0-35 months, exclude AP/HP/MP/TN/WB
 Differences CDF(1998-99)-CDF(1992-93), RURAL

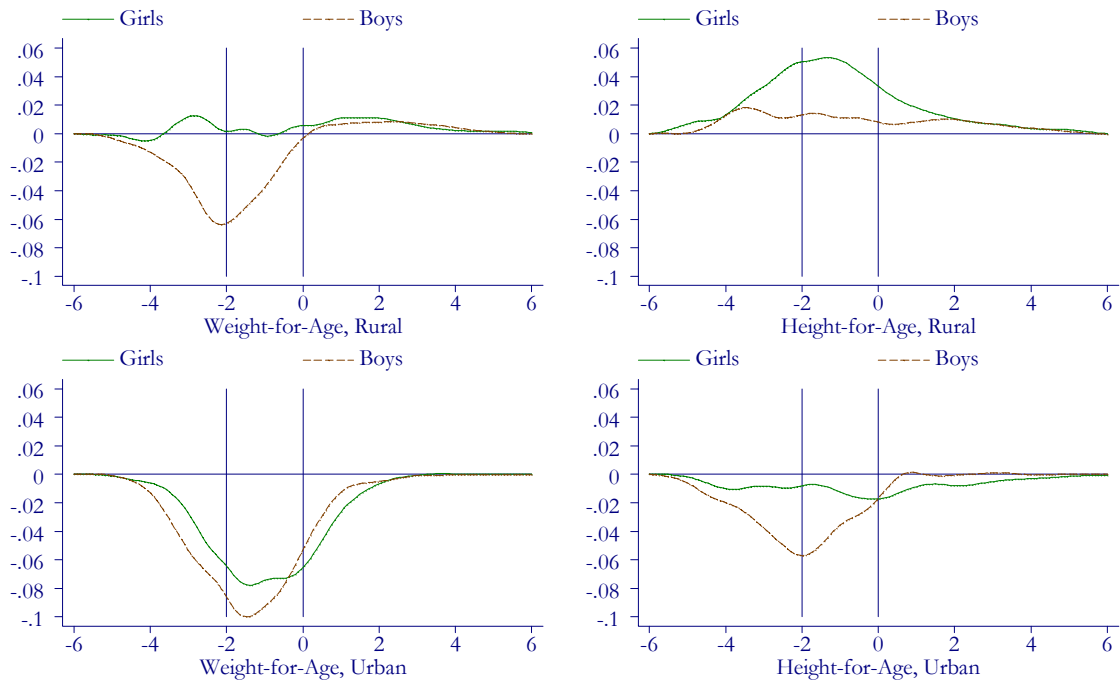


age 0-35 months, exclude AP/HP/MP/TN/WB
 Differences CDF(1998-99)-CDF(1992-93), URBAN

Figure 4 – Change in nutritional status over time. Source: authors' calculations from NFHS-I and NFHS-II. All z-scores are calculated excluding states for which height was not recorded in 1992-93 (Andhra Pradesh, West Bengal, Himachal Pradesh, Madhya Pradesh, and Tamil Nadu).



Age 0-35 months, Guj/Har/Jam/Pun/Raj/UP/Del
Differences CDF(boys)-CDF(girls), North



Age 0-35 months, Guj/Har/Jam/Pun/Raj/UP/Del
Differences CDF(1998-99)-CDF(1992-93) - North

Figure 5a – Boys vs. girls nutritional status and changes over time, North India. Source: authors' calculations from NFHS-I and NFHS-II. North comprises Gujarat, Haryana, Jammu, Punjab, Rajasthan, Uttar Pradesh, and New Delhi. States for which height was not recorded in 1992-93 are excluded.

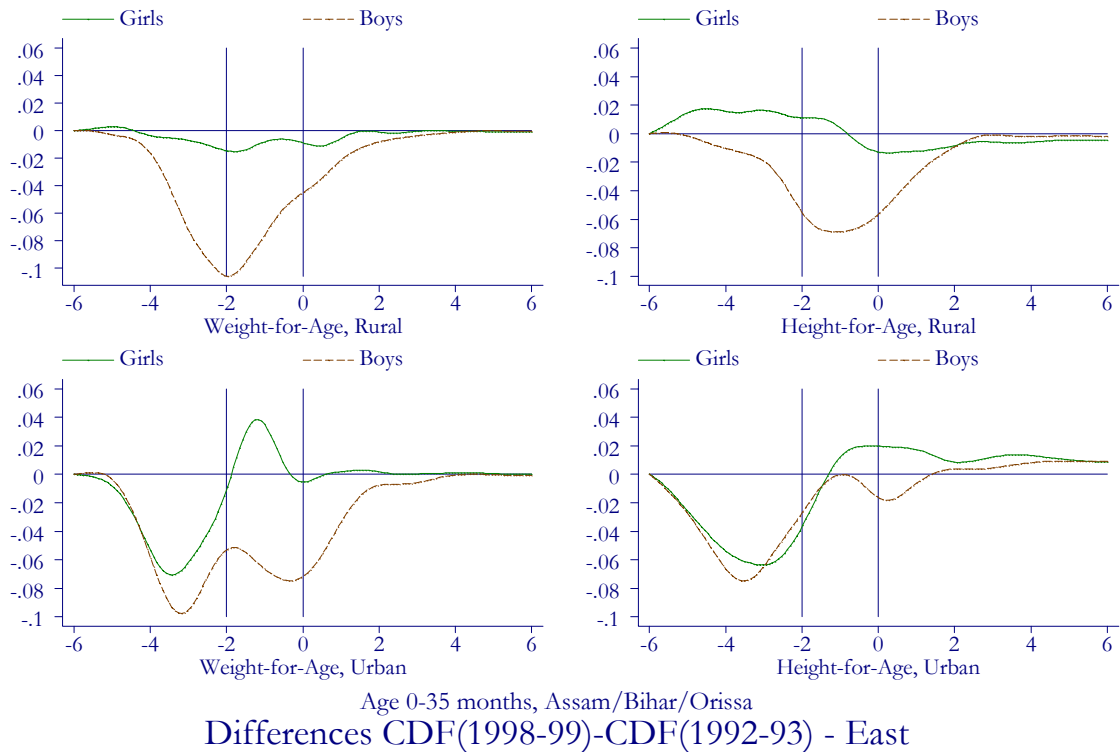
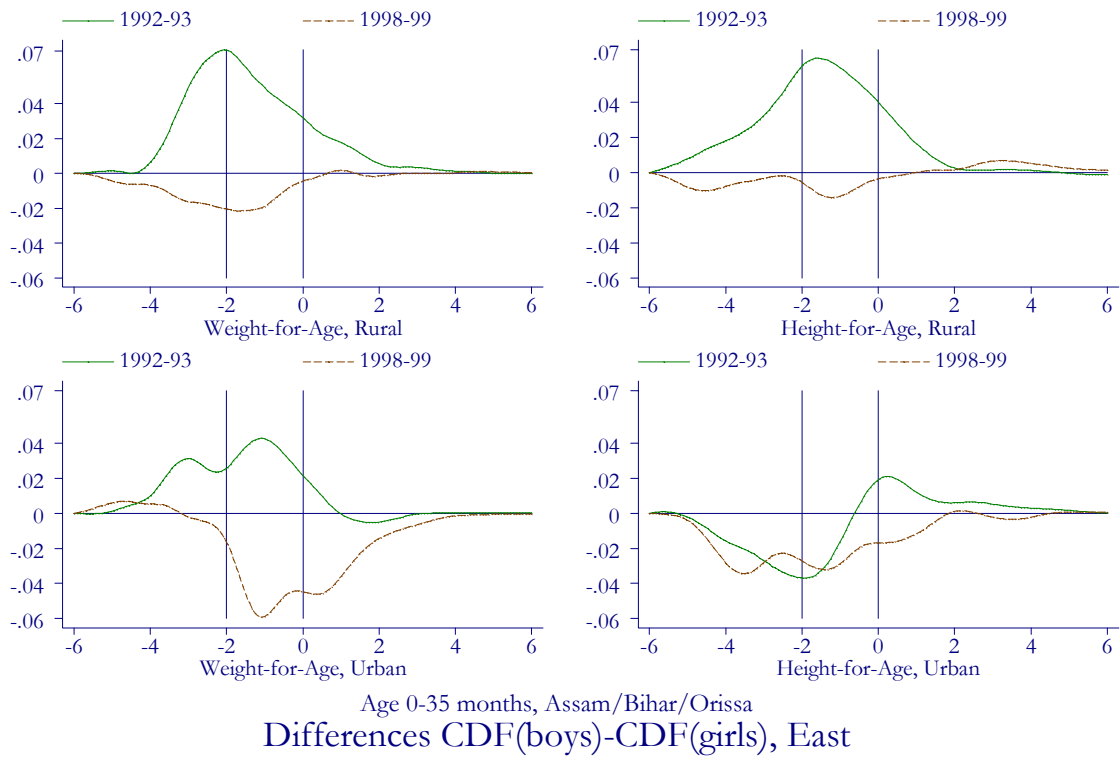
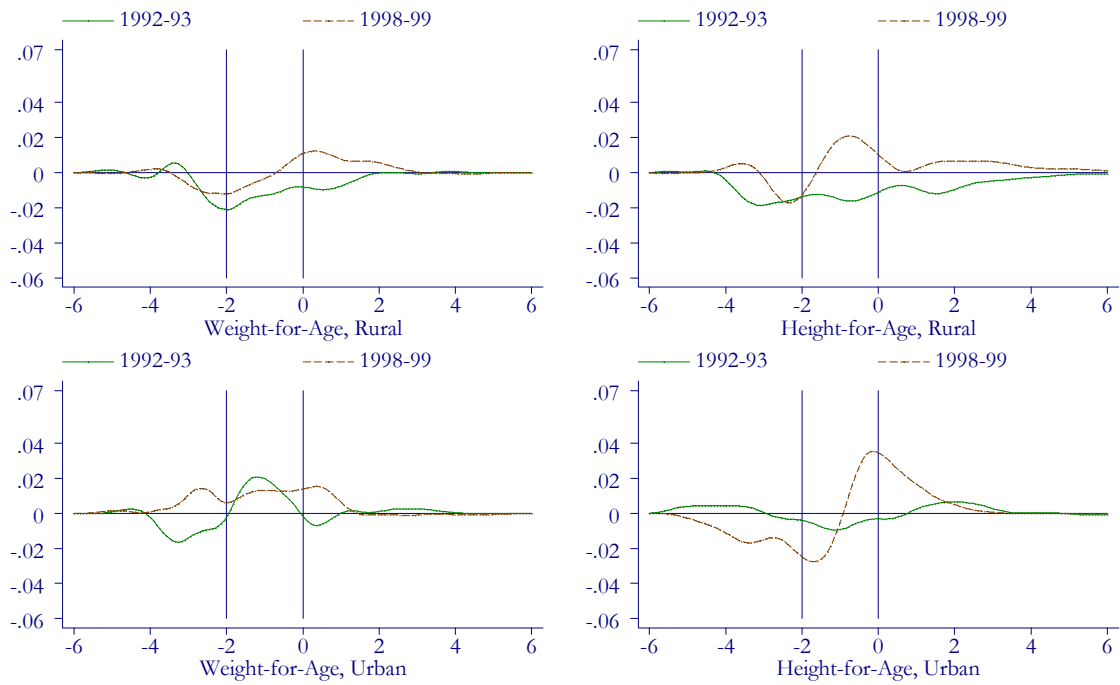
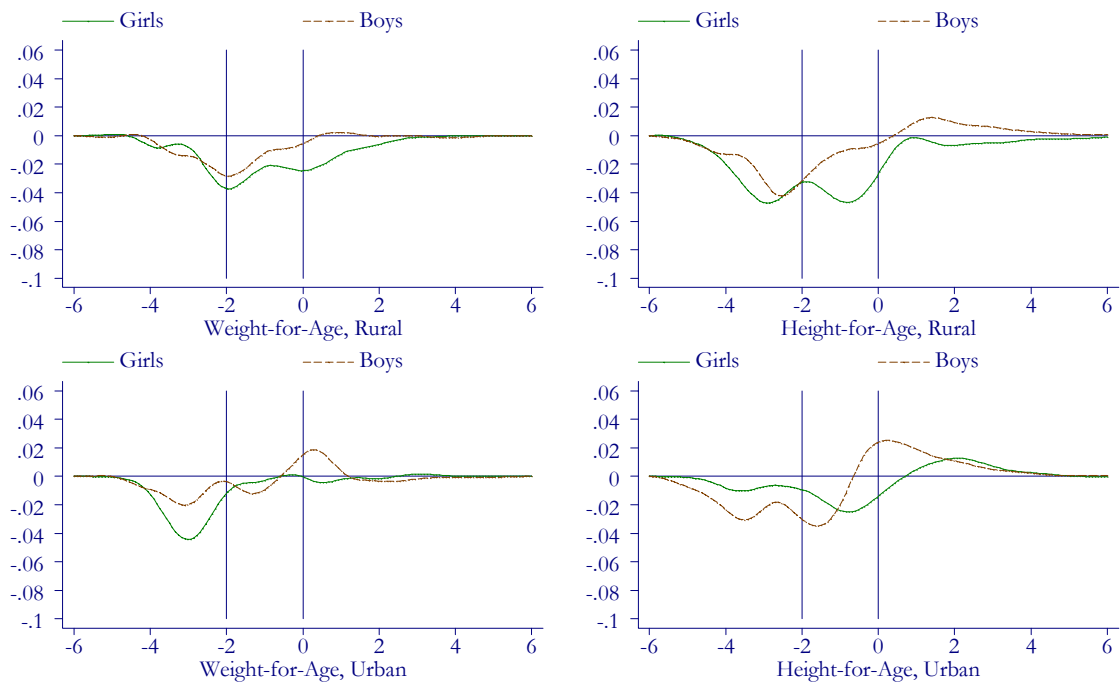


Figure 5b – Boys vs. girls nutritional status and changes over time, East India. Source: authors' calculations from NFHS-I and NFHS-II. East comprises Assam, Bihar, and Orissa. States for which height was not recorded in 1992-93 are excluded.



Age 0-35 months, Kerala/Karnataka/Maharashtra
Differences CDF(boys)-CDF(girls), South



Age 0-35 months, Kerala/Karnataka
Differences CDF(1998-99)-CDF(1992-93) - South

Figure 5c – Boys vs. girls nutritional status and changes over time, South India. Source: authors' calculations from NFHS-I and NFHS-II. South comprises Kerala, Karnataka, and Maharashtra. States for which height was not recorded in 1992-93 are excluded.

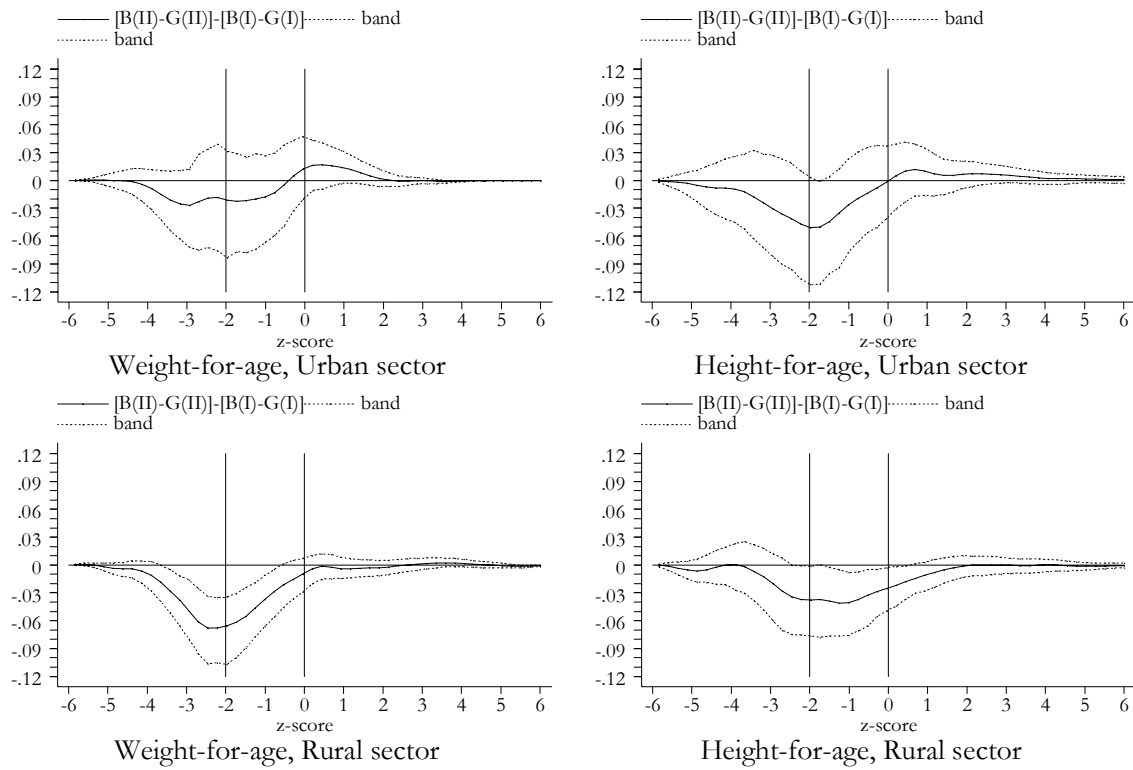


Figure 6a – North. North includes Gujarat, Haryana, Jammu, Punjab, Rajasthan, Uttar Pradesh, and New Delhi.

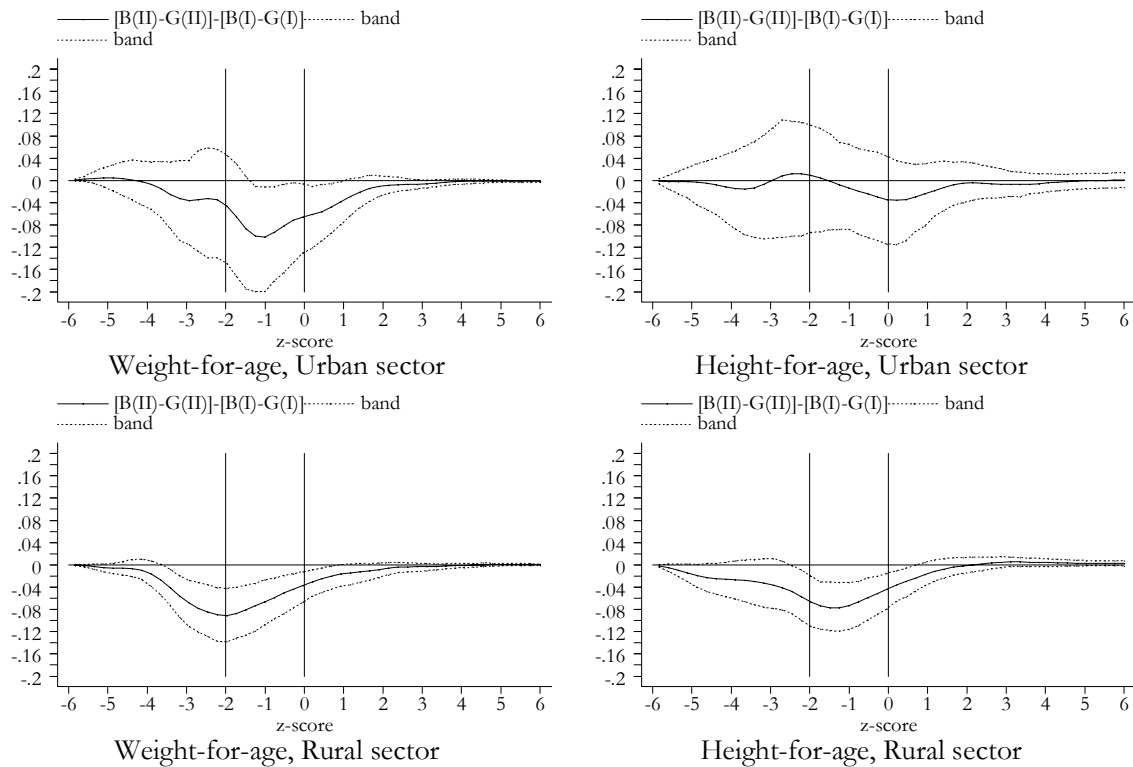


Figure 6b – East. Source: authors' calculations from NFHS-I and NFHS-II. East includes Assam, Bihar, and Orissa.

Source: authors' calculations from NFHS-I and NFHS-II. Children below 3 years old. States for which height was not recorded in 1992-93 are excluded. Each continuous line represents the change over time of the pointwise gender difference in distributions. By construction, negative differences-in-differences (dd) indicate an in(de)crease in boy ad(disad)vantage. The dotted lines represent 95% confidence bands, calculated, at each point over the grid, as the 2.5 and 97.5 percentile of the distribution of estimated dd over 250 bootstrap replications, each done by resampling clusters separately from each survey.

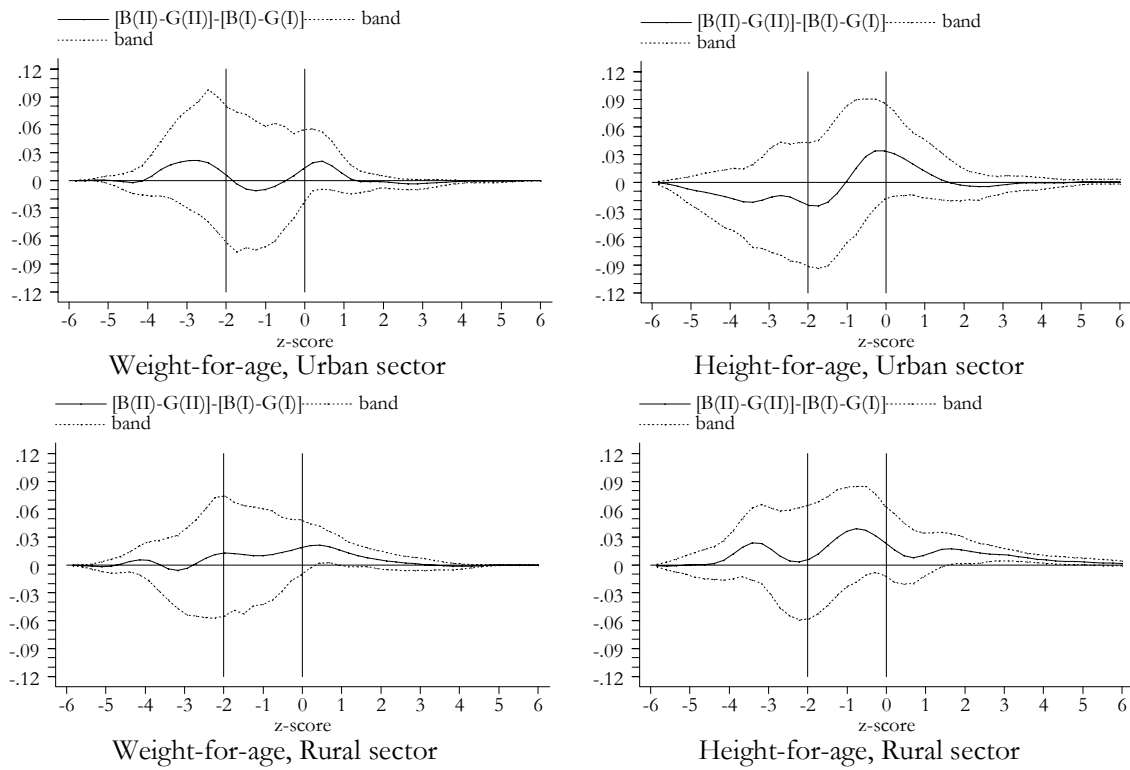


Figure 6c – South. South includes Kerala, Karnataka, and Maharashtra.

Source: authors' calculations from NFHS-I and NFHS-II. Children below 3 years old. States for which height was not recorded in 1992-93 are excluded. Each continuous line represents the change over time of the pointwise gender difference in distributions. By construction, negative differences-in-differences (dd) indicate an in(de)crease in boy ad(disad)vantage. The dotted lines represent 95% confidence bands, calculated, at each point over the grid, as the 2.5 and 97.5 percentile of the distribution of estimated dd over 250 bootstrap replications, each done by resampling clusters separately from each survey.

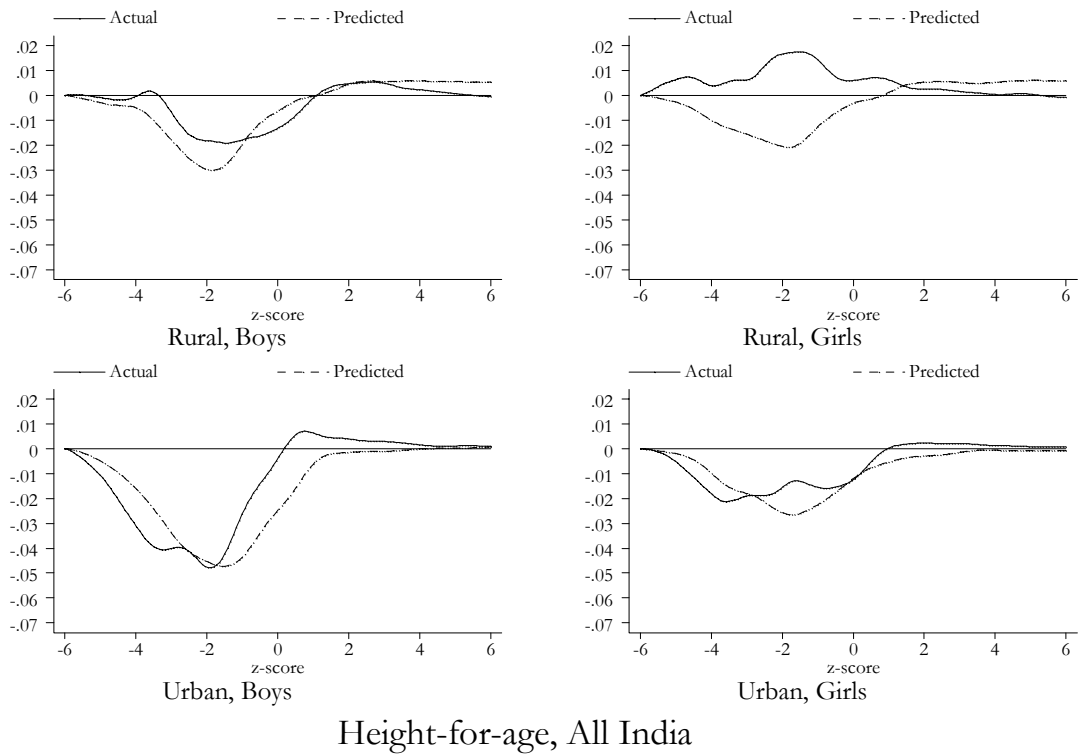
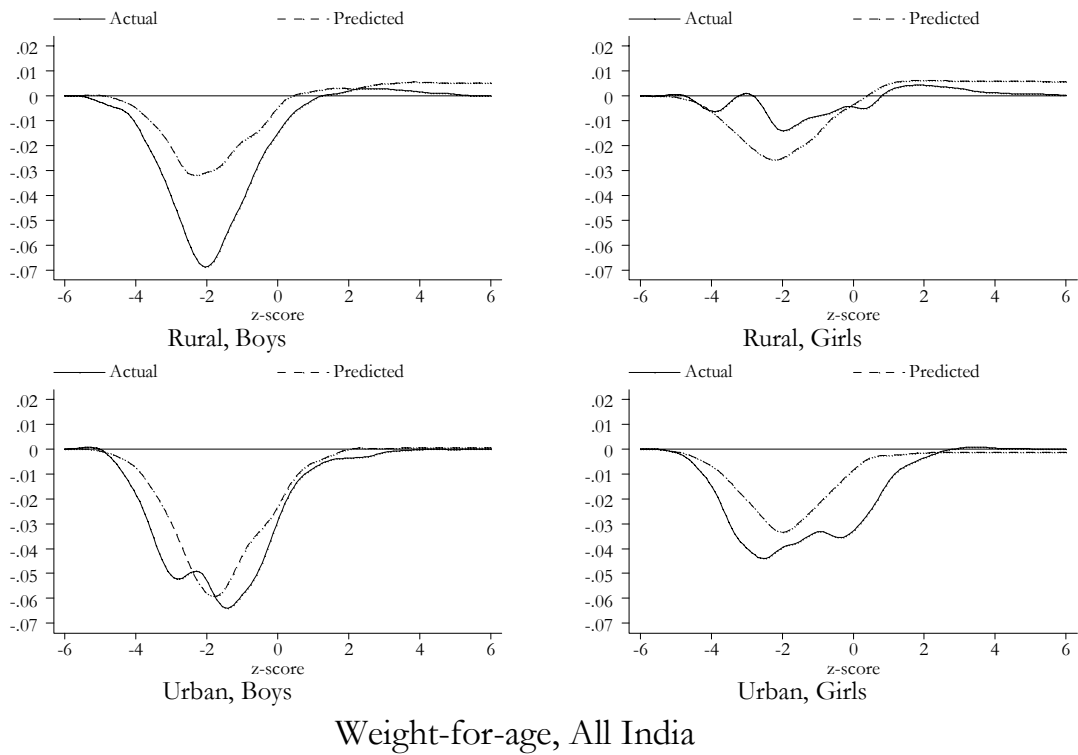


Figure 7 – Actual vs. predicted changes in nutritional status. Source: authors' calculations from NFHS-I and NFHS-II. States for which height was not recorded in 1992-93 are excluded. For details on the construction of the predicted changes see text.

Table 1 - Summary statistics - Women

	1992-93 (NFHS-I)		1998-99 (NFHS-II)					
	Urban	Rural	Urban	Rural				
No. of households	88562		92486					
No. of ever married women age 15-49	89777		90303					
No. of ever married age 13-14	271		0					
% living in rural areas (weighted)	73.8		73.8					
Means (Weighted)	Urban	Rural	Urban	Rural				
Family - Fertility								
Age at first marriage	17.9	16.2	18.2	16.4				
Household size	6.73	7.24	6.48	6.93				
# children below age 5	0.91	1.14	0.81	1.03				
Not using any contraceptive	51.9	65.0	45.5	58.1				
Contraceptive: Female sterilization	28.6	24.9	33.7	31.4				
Contraceptive: Pill	1.8	0.9	2.5	1.8				
Contraceptive: Condom	5.5	1.2	6.8	1.5				
% desiring 3 children or less*	80.3	65.1	85.6	73.1				
% desiring 2 children or less*	56.6	34.4	67.4	46.0				
Desired % of females*	44.1	40.4	45.2	42.2				
North	41.9	38.0	43.3	39.4				
East	43.3	40.4	44.8	42.4				
South	46.3	43.5	47.2	45.6				
Education and Labor Force Participation								
Working	21.1	37.3	24.0	42.0				
North	16.5	29.5	21.2	37.3				
East	16.2	26.7	16.3	27.4				
South	27.5	58.3	29.3	61.2				
Of which, Working for salary	89.1	60.2	89.0	62.6				
North	88.5	43.0	87.2	43.9				
East	88.0	69.7	93.5	79.5				
South	89.8	68.6	89.5	70.4				
	<u>Woman</u>	<u>Partner</u>	<u>Woman</u>	<u>Partner</u>	<u>Woman</u>	<u>Partner</u>	<u>Woman</u>	<u>Partner</u>
% no education	35.6	17.1	71.0	40.5	29.2	13.5	62.4	34.1
North	42.1	18.4	78.4	39.9	33.9	14.0	70.5	32.0
East	37.7	19.9	71.7	43.5	30.3	15.2	64.6	38.9
South	28.5	14.7	59.8	38.7	24.1	12.4	50.0	32.4
% complete secondary or above	10.7	27.0	0.8	8.5	32.8	50.0	7.7	23.2
North	12.1	28.9	0.7	10.2	34.9	53.9	6.1	27.0
East	11.4	30.1	0.7	8.1	29.1	47.5	5.7	19.5
South	9.1	24.0	1.2	6.5	31.9	47.1	11.4	22.0

Source: Author's calculations from NFHS-I and II. All means and proportions calculated using sampling weights.

* Calculated including only numeric answers (this excludes responses such as "up to God" etc.).

North: Delhi, Uttar Pradesh, Rajasthan, Punjab, Jammu, Himachal Pradesh, Madhya Pradesh, Gujarat; East: Assam, Bihar, Orissa, West Bengal; South: Andhra Pradesh, Kerala, Karnataka, Maharashtra, Tamil Nadu. All statistics are calculated including only women of age 15-49.

Table 2 - Summary statistics - Children

	1992-93 (NFHS-I)				1998-99 (NFHS-II)			
	Urban		Rural		Urban		Rural	
Male/Female survival*	-0.0007 (.0031)	0.0107 (.0023)	0.0017 (.0027)	0.0115 (.0021)				
North	.0063 (.0041)	.0249 (.0037)	.0031 (.0039)	.0213 (.0030)				
East	-.0017 (.0091)	.006 (.0047)	.0067 (.0069)	.0062 (.0038)				
South	-.0076 (.005)	-.0052 (.0038)	-.0011 (.0044)	.0034 (.0041)				
<hr/>								
Anthropometric Indicators	Girls		Boys		Girls		Boys	
% with z-score below -2**								
Weight-for-age	44.2 (1.54)	44.3 (1.48)	52.9 (0.83)	55.3 (0.86)	40.6 (1.44)	39.3 (1.29)	51.4 (0.81)	47.8 (0.80)
Obs.	2871	3043	7129	7349	2352	2586	6605	7321
North	46.1 (2.05)	45.6 (2.02)	52.4 (1.07)	53.8 (1.14)	40.0 (2.00)	37.2 (1.70)	52.1 (1.08)	47.1 (1.13)
Obs.	1465	1648	3695	3808	1118	1282	2975	3349
East	47.0 (4.32)	48.4 (3.42)	56.2 (1.82)	63.6 (1.77)	45.4 (3.58)	45.5 (4.21)	54.7 (1.42)	51.8 (1.25)
Obs.	493	478	1471	1499	263	255	1625	1832
South	41.5 (2.68)	41.2 (2.80)	51.2 (1.84)	49.3 (1.63)	40.7 (2.52)	41.1 (2.25)	48.0 (2.16)	45.5 (2.09)
Obs.	563	526	1124	1226	652	679	857	943
Height-for-age	40.0 (1.45)	40.1 (1.49)	48.6 (0.84)	49.8 (0.84)	38.9 (1.41)	35.6 (1.23)	50.7 (0.83)	48.6 (0.79)
Obs.	2871	3043	7129	7349	2352	2586	6605	7321
North	43.9 (1.95)	45.1 (2.03)	50.5 (1.13)	50.0 (1.12)	43.3 (2.02)	39.3 (1.65)	55.4 (1.08)	51.9 (1.08)
Obs.	1465	1648	3695	3808	1118	1282	2975	3349
East	47.7 (4.49)	41.6 (4.02)	50.0 (1.80)	57.6 (1.72)	42.1 (3.76)	38.7 (4.17)	52.6 (1.36)	51.7 (1.30)
Obs.	493	478	1471	1499	263	255	1625	1832
South	32.6 (2.09)	32.3 (2.59)	42.8 (1.76)	39.8 (1.64)	33.0 (2.18)	29.8 (2.09)	40.0 (2.22)	38.3 (1.98)
Obs.	563	526	1124	1226	652	679	857	943
Weight-for-height	16.9 (1.22)	18.6 (0.95)	17.7 (0.64)	21.7 (0.71)	11.5 (0.90)	13.6 (0.88)	16.9 (0.59)	16.6 (0.60)
Obs.	2883	3053	7155	7376	2370	2601	6652	7355
North	16.4 (1.50)	18.1 (1.32)	16.6 (0.80)	19.7 (0.92)	8.2 (0.95)	10.6 (1.12)	12.4 (0.73)	11.8 (0.65)
Obs.	1472	1652	3704	3821	1131	1295	3002	3370
East	11.8 (2.42)	19.5 (2.66)	18.7 (1.41)	25.7 (1.62)	15.9 (2.39)	20.7 (3.15)	20.9 (1.10)	21.6 (1.11)
Obs.	493	479	1476	1501	264	256	1629	1833
South	19.7 (2.50)	19.5 (1.69)	19.7 (1.55)	21.9 (1.27)	14.8 (1.71)	16.3 (1.46)	21.9 (1.46)	21.0 (1.63)
Obs.	568	530	1133	1231	656	679	868	950

Source: Author's calculations from NFHS-I and II. Calculations done with sampling weights. Standard errors in parenthesis.

*Calculated as (sons alive/sons ever born) - (daughters alive/daughters ever born). By construction, this is calculated only for women who had at least one birth of both sexes. The standard errors take into account clustering. The figures in bold indicate that the null hypothesis of mean equal to zero is rejected using a 5% significance level.

**All results are calculated for children up to 3 years old living in states where height was recorded in NFHS-I. The remaining states are categorized as follows. North: Delhi, Uttar Pradesh, Rajasthan, Punjab, Jammu, Gujarat; East: Assam, Bihar and Orissa; South: Kerala, Karnataka, and Maharashtra. The results for all India also include union territories.

Table 3 - Tests of comparisons of distributions

	$H_0: CDF_{\text{boys}} = CDF_{\text{girls}}$				$H_0: CDF_{92/93} = CDF_{98/99}$			
	1992-93		1998-99		Boys		Girls	
	Rural	Urban	Rural	Urban	Rural	Urban	Rural	Urban
Weight-for-age	0.1189**	0.3090	0.0044##	0.6874	0.0000***	0.0000***	0.0429***	0.005***
North	0.6475	0.4927	0.0012###	0.1491	0.0001***	0.0002***	0.0272	0.0022***
East	0.0079***	0.4937	0.7233	0.6628##	0.0000***	0.0333**	0.8874	0.0733
South	0.7864	0.2962	0.6336	0.8582	0.8560	0.2488	0.2025	0.6193
Height-for-age	0.1862**	0.7704	0.2535	0.3786	0.0588**	0.0151***	0.2632	0.8001
North	0.0629	0.9162	0.0385###	0.5041	0.2465#	0.3978#	0.0004###	0.9571
East	0.0644***	0.1978	0.9031	0.7817**	0.0033***	0.0945	0.6964	0.4571
South	0.3412	0.9646	0.4968	0.6361	0.0362	0.4166	0.0580	0.9265
Weight-for-height	0.0000**	0.0841***	0.3731	0.1487	0.0000***	0.0036***	0.0180***	0.0051**
North	0.0453	0.0504*	0.3432	0.0336	0.0000***	0.0000***	0.0007***	0.0000***
East	0.0259***	0.3515**	0.1851	0.5349#	0.0001***	0.2578**	0.0006**	0.4564
South	0.1164	0.7653	0.8001	0.2919	0.4157	0.4971	0.8018	0.1001

Source: Author's calculations from NFHS-I and II. The p-values refer to test of the null hypothesis indicated in the column header. All these tests are adjusted for the complex survey design. For details on the construction of the tests see Section 4. Figures in bold represent p-values below .10. The "*" reported in the table refer to Kolmogorov-Smirnov type test of first-order stochastic dominance where, for each pair of distributions, the null hypothesis is that the wave(gender)-specific distribution for boys (wave I) is everywhere *not above* the corresponding distribution for girls (wave II). So, a "*" indicates female advantage in the first four columns, while it indicates improvement over time in the last four columns. * indicates rejection using a 10% significance level, ** 5% level, *** 1% level. The "#" reported in the table indicate rejections for the same kind of tests where the role of the two distributions in the null and alternative is reversed. So "#" denotes male advantage in the first four columns, and worsening over time in the last four columns.

These tests for first-order stochastic dominance are not adjusted for the complex survey design, so that the positive intracluster correlation observed in the sample typically implies that the tests will not have correct size.

All results are calculated for children up to 3 years old living in states where height was recorded in NFHS-I. The included states are categorized as follows. North: Delhi, Uttar Pradesh, Rajasthan, Punjab, Jammu, Gujarat; East: Assam, Bihar and Orissa; South: Kerala, Karnataka, and Maharashtra. The results for all India also include union territories.