Child Nutrition in India in the Nineties

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I. Introduction

India experienced several years of fast economic growth during the 1990s, and according to many observers this period also saw a considerable decline in poverty, especially in urban areas (see, e.g., Deaton and Drèze 2002; Deaton 2003; Tarozzi, forthcoming).[1] The first objective of this article is to document the extent to which the 1990s saw a reduction in malnutrition among very young children (those less than 3 years old); second, we study whether changes in child growth performance have been similar for boys and girls and in different geographical areas; third, we provide a first attempt at explaining the observed trends. The source of our data is the Indian National Family and Health Survey (NFHS), a data set that contains detailed information on health and fertility for two independent cross sections of ever-married women of fertility age, the first from 1992–93 and the second from 1998–99.

Many researchers have documented the presence in India of widespread child malnutrition, as measured by anthropometric indicators such as weight or height (e.g., Klasen 1999; Svedberg 2000). The reduction of child mal-

For useful comments, we would like to thank Orazio Attanasio, Sonia Bhalotra, Angus Deaton, William Dow, Jean Drèze, Gayatri Koolwal, David McKenzie, Dilip Mookherjee, John Strauss, Duncan Thomas, two referees, and seminar participants at University of California, Berkeley, Boston University, Princeton University, RAND, University of California, Los Angeles, the 2005 North East Universities Development Consortium conference (Brown University), the workshops "Indian Economy: Policy and Performance, 1980–2000" (University of British Columbia) and "Human Development in India: Microdata Perspectives" (New Delhi), the Pacific Conference for Development Economics (Santa Clara University), and the Third Minnesota International Economic Development Conference. We are very grateful to Macro International, Inc., for granting us access to the data. Last but not least, Joanne Yoong provided excellent research assistance. We are solely responsible for all errors and for all views expressed in this article.

[1] For a broad overview of the debate on poverty reduction in India over this period, see the collected essays in Deaton and Kozel (2005), which also include less optimistic assessments of the degree of poverty reduction, as in Datt, Kozel, and Ravallion (2003) or Sen and Himanshu (2004).

 0013-0079/2007/5503-0004$10.00

nutrition is certainly one of the most desirable components of economic development. Not only is child malnutrition strongly associated with increased child mortality and morbidity but there is now ample evidence that inadequate nutrition in childhood (and in utero) hinders long-term physical development, reduces the development of cognitive skills, and, as a consequence, affects negatively schooling attainment and several outcomes later in life, including productivity, mortality, and the likelihood of developing chronic diseases (see Strauss and Thomas [1998], Behrman, Alderman, and Hoddinott [2004], and Maluccio et al. [2006] for extensive references).

The analysis of gender differences plays a very important role in our analysis. Preference for sons over daughters and gender inequality are well-known and still widespread realities in India, particularly in the North-West; they are reflected in such phenomena as sex-selective abortion and female disadvantage along such crucial dimensions as schooling, health and health care, and child mortality. Some studies also find gender differences in nutrient intakes and nutritional status (Behrman 1988a, 1988b), although these findings are not confirmed in other studies, as discussed in Harriss (1995). Several studies from such different disciplines as anthropology, economics, and sociology have found that the preference for sons is particularly strong in those areas where the cultural, social, and economic roles of women in society and/or within the household are weaker because, for instance, women are less important as bread earners, dowries are more common, or bequests favor sons over daughters (see, e.g., Miller [1981], Basu [1992], Dasgupta [1993], Murthi, Guio, and Drèze [1995], and Drèze and Sen [2002] for extensive references).

Many of the factors associated with gender inequality appear to be related to the presence of economic constraints.[2] In a seminal paper, Rosenzweig and Schultz (1982) suggest that the preferential treatment of boys may be an unfortunate but rational response to unequal economic "returns" to boys and girls and hence can coexist with the absence of differences in the way the welfare of boys and that of girls enter the parents' utility function. The authors use this argument to explain the correlation between gender bias in survival rates and female labor market participation. Behrman (1988a, 1988b), using data from a small number of Indian villages, finds that parents favor equal treatment of children, but he also finds evidence of promale bias in the intrahousehold allocation of resources during the lean season, a time when resource constraints are more likely to bind. Jensen (2002), building on insights from Yamaguchi (1989), discusses how gender bias in average outcomes may arise

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

even if females are not discriminated against in the intrahousehold allocation of resources if girls are more likely than boys to live in families with more siblings and hence fewer resources per head. Such differences in the number of siblings may emerge if preference for sons induces families to have more children when they have not yet achieved the desired number of boys.

The fact that resource constraints—coupled with promale bias in economic opportunities—appear to provide an economic "rationale" for the existence of gender bias might lead one to expect a path toward 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 in which 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 such elements 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. When evaluating gender differences, another advantage of nutritional status versus, say, nutrient intakes, morbidity, or health care is that the former is relatively easily measured and therefore much less prone to measurement error or reporting bias.

To evaluate changes in nutritional status, we transform the anthropometric indicators into z-scores, that is, we normalize the indicators by using the mean and the standard deviation of the same index for children of the same gender in a reference population. The use of z-scores is common in nutritional studies (more on this below), and two reasons make its use 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 where boys and girls are, on average, equally well nourished. Second, it allows us 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 cumulative distribution functions of z-scores between genders (for a given wave) or over time (for a given gender).

Overall, we find that, in urban areas, child nutritional status in India improved substantially during the 1990s. In rural India, which accounts for the bulk of the total population, our results show large improvements in short-term measures of nutritional status, while height-for-age (a measure of long term nutritional status) improved much less. We also find that gender inequality in nutritional status increased, with nutritional status improving substantially more for boys than for girls. We also document the existence of apparent geographical differences in these changes: the gender differences in the changes in nutritional status are particularly striking in rural areas of North India and East India, areas where the existence of widespread son preference has been documented by an immense body of research.

In the second part of this article, we explore alternative explanations for the observed trends. First, we consider (and exclude) the possibility that rural-to-urban migration and changes in infant mortality are driving the differences in the changes between sectors and genders. Second, we study the relation between changes in height-for-age and changes over time in a list of child-, household-, and community-specific economic and demographic variables that should be strongly associated with child growth performance. Overall, we find that the sole change in the distribution of the predictors leaves much of the actual changes unexplained, especially in rural areas. Oaxaca decompositions of the probability of stunting confirm that, for both genders and across all of India, most of the change in anthropometric performances is explained by changes in the regression coefficients that relate the z-scores to the predictors rather than by changes in the predictors themselves. However, a detailed analysis of the patterns of the changes in the coefficients does not point to a simple explanations for the emerging gender differences we document.

This article proceeds as follows. In the next section, we describe the data set. In Section III, we discuss the anthropometric indicators that represent the main outcome of interest of our analysis. In Section IV, 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 V, we provide a first attempt at explaining the observed changes. Section VI concludes.

II. Data

The primary source of our data is the two waves of the Indian National Family and Health Survey (NFHS) that were available when this article was being written. The NFHS is one of many demographic and health surveys that have been carried out in several developing countries with the primary purpose of collecting information on health, fertility, and other family issues from ever-married women of fertility age. The first wave of the survey (NFHS-I) was completed between April 1992 and August 1993 with a sample of ever-married women ages 13–49. The second wave of the survey (NFHS-II) was completed between November 1998 and December 1999 and sampled ever-married women ages 15–49.[3] Each survey contains reports from approximately 90,000 women, sampled from all Indian states using a stratified and clustered survey design. In all our calculations, we make use of the sampling weights contained in the survey, and we report separate results for urban and rural areas.

The largest component of the surveys is an individual questionnaire administered to each ever-married woman of fertility age in the sample. The questionnaire 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.[4] In particular, height and weight were measured for children below age 4 in NFHS-I and below age 3 in NFHS-II. Because of the 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. To enhance comparability, we will then 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. A separate questionnaire administered at the household level contains several household characteristics, including a complete household roster, and individual information on work status, educational attainment, and a few selected health indicators. Finally, in rural areas, a village questionnaire records information on village characteristics.

Tables 1 and 2 report selected summary statistics at the household and individual level. For several statistics, we also present a geographical breakdown

[3] We ignore the difference in the lower bound of mother's age in the waves, as fewer than .5% of women in NFHS-I were 13 or 14 years old.

[4] NFHS-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's 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).

TABLE 1
SUMMARY STATISTICS, WOMEN, PART 1

	NFHS-I (1992–93)		NFHS-II (1998–99)	
Number of households	88,562		92,486	
Number of ever-married women ages 15–49	89,506		90,303	
Number of ever-married women ages 13–14	271		0	
% living in rural areas (weighted)	73.8		73.8	
	Urban	Rural	Urban	Rural
Number of children ages 0–35 months	9,357	24,826	7,609	21,053
Number of children ages 36-47 months	3,080	8,012	0	0
Means (weighted):				
Age at first marriage	17.9	16.2	18.2	16.4
Household size	6.73	7.24	6.48	6.93
Number of children below age 5	.91	1.14	.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	.9	2.5	1.8
Contraceptive: condom	5.5	1.2	6.8	1.5
% desiring three children or less*	80.3	65.1	85.6	73.1
% desiring two children or less*	56.6	34.4	67.4	46.0

Source. Authors' calculations.
* Calculated including only numeric answers (while excluding responses such as "up to God," etc.). All statistics are calculated including only women of ages 15–49.

following the geocultural classification proposed by Sopher 1980 and used by, among others, Dyson and Moore (1983), Bourne and Walker (1991), and Dasgupta (1993). The major Indian states are then grouped into three regions 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. All results reported by region in this article only include the major Indian states listed above, while they exclude the Union Territories (which account for less than 5% of the population).

Many indicators suggest that important changes are taking place. The figures in table 1 show a fertility decline in both urban and rural areas. Average household size declined by about .2 persons, and the number of children below age 5 decreased by about .1 persons.[5] We also observe a decline in desired family size, defined as the ideal number of children that a respondent with no children would like to have or the ideal number that a woman with children would have liked to have had if she could go back to the time when she did not have any children. The use of contraceptives increases between the two

[5] However, some observers, citing evidence from other data sources, have suggested that NFHS-II may have underreported the number of births (see Irudaya Rayan and James [2004] and references therein).

TABLE 2
SUMMARY STATISTICS, WOMEN, PART 2

	NFHS-I (1992–93)		NFHS-II (1998–99)	
	Urban	Rural	Urban	Rural
Number of children ages 0–35 months	9,357	24,826	7,266	21,053
North	4,117	10,870	3,371	9,510
East	1,251	4,424	963	4,779
South	1,797	3,969	2,209	3,351
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
% women 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
% women working who receive earnings	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
% illiterate	36.8	72.4	33.2	66.9
North	42.6	79.4	36.1	74.2
East	36.7	72.1	32.9	67.8
South	31.7	63.5	29.9	57.7
% women completed secondary or above	28.5	5.3	32.8	7.7
North	29.4	4.6	34.9	6.1
East	25.9	4.2	29.1	5.7
South	28.7	7.1	31.9	11.4
% partners with no education	17.1	40.6	13.7	34.6
North	18.5	40.1	14.5	32.7
East	20.0	43.8	15.7	39.8
South	14.8	38.8	12.5	32.6
% partners completed secondary school or above†	27.1	8.5	30.7	10.5
North	29.0	10.2	35.2	12.7
East	30.2	8.1	31.8	8.9
South	24.1	6.5	26.0	9.0

Source. Authors' calculations.
Note. Statistics reported to the right of the variable description refer to all India. All statistics are calculated including only women of ages 15–49.
* Calculated including only numeric answers (while excluding responses such as "up to God," etc.).
† Binary variable equal to one for men with at least 12 completed years of education.

surveys. In urban areas, the proportion of women who do not practice any form of birth control declined from 52% to 45%. In rural areas, the proportion declined from 65% to 58%.

The figures reported in table 2 refer to variables that have often been used as indicators of gender inequality. These include direct measures of preference for sons, male versus female schooling achievement, and the role of women as bread earners. The desired proportion of girls, calculated from numerical answers to direct questions about the "ideal" number or sons and daughters, displays the expected North-South gradient, with much stronger son preference in the North, especially in rural areas. It is interesting that, in every region and sector, the mean proportion of children that is desired to be girls is higher in 1998–99 than in 1992–93, even if all the figures remain below one-half. In the North, the proportion of desired girls increases by approximately 1 percentage point 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% in cities and from 43.5% to 45.6% in rural areas. Similar patterns emerge in Eastern states.[6]

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 areas than in urban areas; participation rates in urban areas 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% in urban areas and from 29.5% to 37.3% in rural areas. Third, participation rates are about twice as large in the South as compared to the North, both in cities and in villages. For example, in NFHS-II, 61.2% of women worked in the rural South, while only 37.3% did in the rural North. In eastern states, women's work participation is even lower than in the North. The proportion of working women who are also earning money shows instead a very stable picture. In urban areas, the fraction remains close to 90% in all regions. While in the rural South, approximately three-quarters of working women also receive earnings, the proportion is only two-thirds as large in the North.

Female illiteracy rates once again confirm the familiar North-South pattern. In 1992–93, almost 80% of ever-married women of fertility age who lived in rural areas of northern states had no formal education. In the South, the proportion was still very high, but it was about 15 percentage points lower as compared to the North. The gradient is also clearly present in urban areas, where illiteracy rates are, however, approximately only half as large. Illiteracy is significantly less common among the partners of the women in the sample. 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 general attitudes toward) 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 for both men and women. In urban areas, the percentage of women with

[6] If we interpret nonnumerical 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.

at least a secondary degree increases from 28.5 in NFHS-I to 32.8 in NFHS-II. The figures are much lower in rural areas, but in relative terms the increase is larger, as the overall fraction increases from 5.3% to 7.7%. These statistics are clearly very rough measures of the socioeconomic status of women in India, but overall they seem to suggest that it somehow improved during the 1990s.

In the next section, we turn to the description of the anthropometric indicators of child nutrition that form the core of our analysis. The first rows of table 2 show that the number of children ages 0–3 in both NFHS-I and NFHS-II is quite large, even when we disaggregate at the sector and region level; that number ranged from 963 in the urban East in 1992–93 to 10,870 in the rural North in 1998–99.

III. Child Nutritional Status: Measurement

The use of anthropometric indices to evaluate child nutritional status is a well-established practice (see, e.g., Waterlow et al. 1977; WHO Working Group 1986; Gorstein et al. 1994). Height (given age) is the preferred measure of long-term nutritional status as it reflects both current and past nutritional status. 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-height is a better measure of short-term nutritional status. Weight-for-age can also change rapidly, but—unlike weight given height—it does not distinguish between small but well-fed children and tall but thin ones, and so it can be seen as a combination of the other two indices. For this reason, in most of our empirical results, we will omit weight-for-age from the analysis.

Let x_{ig} represent weight or height of a specific child i in a group g. When the indicator measures height, the group is defined by age and gender. When the indicator measures weight, the reference group is identified by gender and either age (in the case of weight-for-age) or height (in the case of weight-for-height). To gauge the nutritional status of a child, it is necessary to compare the child's outcome to a corresponding "normal" outcome for a child that belongs to the same group. The common practice is to make use of z-scores, calculated as $(x_{ig} - x_g)/\sigma_g$, where 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 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% of boys in the reference population with the same height.[7] Children are said to be "stunted" if their height-for-age z-score

[7] In reality, anthropometric indicators are not exactly described by normal distributions. Recently

is below −2 and to be "wasted" if their weight-for-height is below the same threshold.

Both NFHS waves 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). Agarwal et al. (1991) and Bhandari et al. (2002) show 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 chose 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 that is comparable across all age groups and that has a distribution that can be easily tracked over time.

IV. Child Nutritional Status in the 1990s

In figure 1, we plot nonparametric locally weighted regressions (Fan 1992) of z-scores on age, pooling all observations from NFHS-I. All the patterns of the z-scores are consistent with what is commonly observed in low-income countries (see, e.g., Shrimpton et al. 2001) and show weight and height performances that are, on average, well below those of the American children in the reference population. The curve for weight-for-age starts below zero, declines until the age of about 18 months, and then stabilizes below −2. The mean weight performance is therefore approximately equal to that of the first percentile of the reference population. Height-for-age, which represents a measure of long-term nutritional status, presents an even more striking pattern, and the regression is still sloping downward (and approximately equal to −3) for 4-year-old children. Because most low-weight children are also small, the weight-for-height indices show a degree of wasting much lower than that of

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).

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

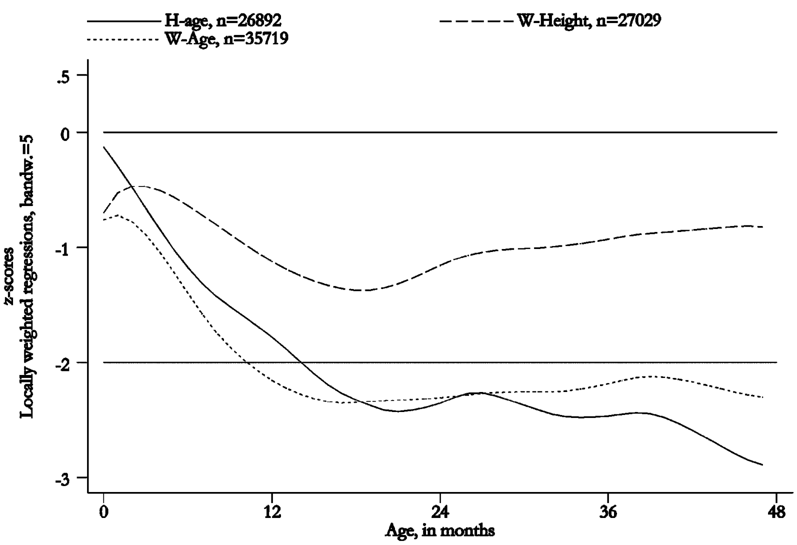

Figure 1. Source: authors' calculations from NFHS-I (1992–93). The *z*-scores for weight-for-age are for all Indian states, rural and urban areas, while the *z*-scores for height-for-age and weight-for-height exclude Phase I states (Andhra Pradesh, West Bengal, Himachal Pradesh, Madhya Pradesh, and Tamil Nadu), for which height is missing.

stunting, and the *z*-scores curve remains close to -1. Note also that the degree of wasting decreases for older children.

In order to analyze changes over time in child nutritional status, we study the changes in the whole distribution of *z*-scores for a given geographical area and demographic group. First, we estimate the densities nonparametrically using a biweight kernel and choosing the bandwidth according to the criterion proposed by Silverman (1986). Then, we calculate the cumulative distribution functions by numerically integrating the densities.[9] Finally, for a given value z of the *z*-scores, letting F denote the cumulative distribution function, we calculate the differences in the distributions as $F_{98-99}(z) - F_{92-93}(z)$, so that improvements will be reflected by negative numbers. Other researchers have used analogous differences to evaluate changes over time or discrepancies across countries in the distribution of indicators of nutritional status (see, e.g., Sahn and Stifel 2002; Strauss et al. 2004).

Figure 2 plots the results for weight given height, by sector, for all children less than 36 months old in all Indian states where height was recorded in NFHS-I. In both rural and urban areas, the distributions of *z*-scores shift markedly to the right, indicating large improvements. In the top rows of table

[9] We prefer this estimation strategy to the alternative of estimating cumulative distribution functions directly as the direct estimation of cumulative distribution functions leads to lines that are excessively jagged.

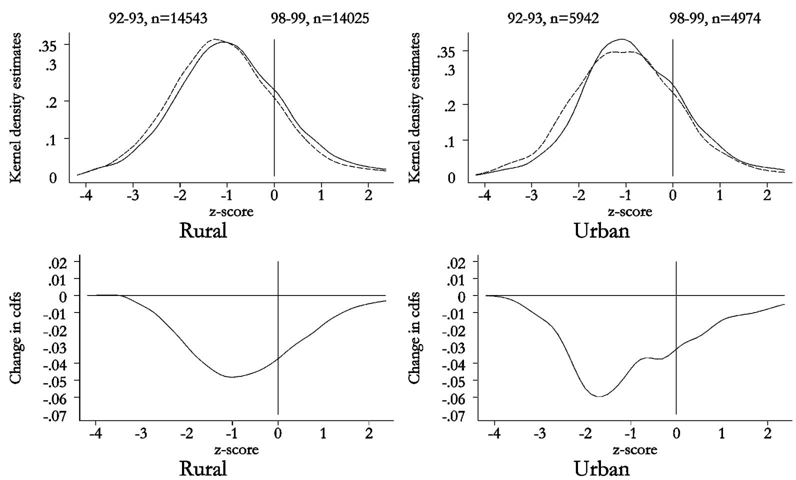

Figure 2. Weight-for-height, both genders, ages 0–3, no Phase I states

3, we report the results of a battery of tests of comparisons between distributions of weight-for-height z-scores. The figures in columns 1 and 4 are p-values of Kolmogorov-Smirnov tests of equality between the two distributions, so that the null hypothesis that is being tested is $F_{98-99}(z) = F_{92-93}(z)\,\forall\, z$. The p-values are calculated using simulations, using the bootstrap procedure described in Abadie (2002). The test statistic is based on the supremum of the absolute value of the differences $F_{98-99}(z) - F_{92-93}(z)$ calculated over a grid of points over the support of z (the difference is rescaled by a factor that is a function

TABLE 3
TESTS OF EQUALITY AND STOCHASTIC DOMINANCE: CHANGES OVER TIME, ALL INDIA

	Rural			Urban		
	p-Value			*p*-Value		
	Equality (1)	FOSD (2)	No FOSD (3)	Equality (4)	FOSD (5)	No FOSD (6)
Weight-for-height	0	1	[−3 1]	0	1	[−3 1]
			[−3 1]			[−3 1]
			NOSD			[−3 −.96]
Height-for-age	.828	.464	NOSD	.06	.764	[−4 −1.55]
			NOSD			[−4 −3.02]
			NOSD			[−4 −3.27]
Weight-for-age	0	.828	[−4 .41]	.004	1	[−4 2]
			[−4 −3.39]			[−4 1.88]
			NOSD			[−4 .78]

Sources. Authors' calculations from NFHS-I and NFHS-II, for all India excluding Phase I states.
Note. Columns 1 and 4 report p-values of Kolmogorov-Smirnov tests of equality of distributions. In cols. 2 and 5, the null is that the distribution in 1998–99 (weakly) first-order stochastically dominates (FOSD) the distribution in 1992–93. Columns 3 and 6 report the results of an intersection-union test of the null of no stochastic dominance (NOSD), using (top to bottom) 10%, 5%, and 1% significance level. See the text for details on the tests. All tests are robust to the presence of intracluster correlation.

of sample size in the two distributions). We use a 50-point grid over the interval [−3 1], and for each test we use 250 replications, adopting block bootstrap to take into account the clustered survey design. In columns 2 and 5, we calculate analogous simulation-based tests for the null hypothesis $F_{98-99}(z) \leq F_{92-93}(z)\ \forall\, z$; that is, we test the null hypothesis that the distribution of z-scores in 1998–99 (weakly) first-order stochastically dominates the distribution in 1992–93. This test statistic is based on the rescaled supremum of the differences described above. In both rural and urban areas, the null of equality is clearly rejected at standard significance levels, while there is strong support for the null of first-order stochastic dominance. Columns 3 and 6 report the results of an intersection-union test of no stochastic dominance; that is, the null hypotheis is $F_{98-99}(z) > F_{92-93}(z)$ for some $z \leq \bar{z}$ (Howes 1996; Davidson and Duclos 2000). The null is rejected in favor of the alternative that the distribution in 1998–99 first-order stochastically dominates the distribution in 1992–93 (or vice versa) if all differences $F_{98-99}(z) - F_{92-93}(z)$ calculated over a grid of points are negative (positive), and all pointwise tests of equality reject the null. For each anthropometric index, we display the range over which the null is rejected using a 10%, 5%, or 1% significance level. We use the same grid as for the Kolmogorov-Smirnov tests, and we calculate all pointwise tests taking into account the presence of intracluster correlation.[10] In urban areas, and using either a 10% or 5% significance level, the null of no stochastic dominance is rejected over the whole grid in favor of the alternative that the more recent distribution first order stochastically dominates the earlier one. In rural areas, the null is rejected over the whole range using a 10% level, and until −.877 using a 5% level. The null is not rejected in either sector if we use a 1% significance level.

The change in the distribution is not only statistically significant but also very large in practical terms. For instance, in rural areas, the proportion of children who are wasted (i.e., whose weight-for-age z-score is below −2) decreased by about 3 percentage points, while the decrease was larger than 5 percentage points in urban areas. These are large changes, especially once we take into account that the two surveys are separated by only 6 years. The changes become even more impressive once we transform these percentages into actual headcounts. According to NFHS-II, the urban sector of Phase II states (for which height in the previous round was recorded) accounted for 18.2% of the total Indian population, and the rural sector accounted for

[10] Note that this test, being an intersection-union test, is quite conservative, so that the actual size will be generally lower than the nominal size. Note also that this test never rejects the null if the chosen grid includes points too far along the tails of the distributions as all cumulative distribution functions are identical (either one or zero) at extreme points.

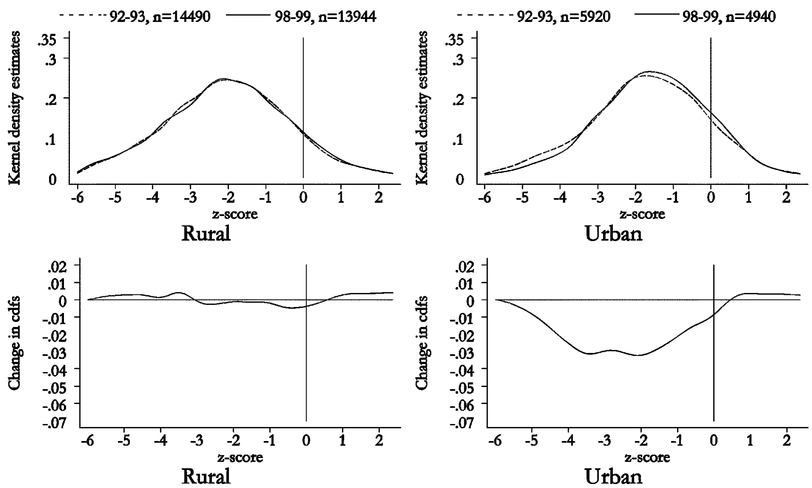

Figure 3. Height-for-age, both genders, ages 0–3, no Phase I states

approximately half. In the same states, children below age 3 represented approximately 6% of the total population in urban areas and 7.5% in the countryside. With the total Indian population reaching one billion at the end of the 1990s, the estimated changes in cumulative distribution functions in urban areas indicate that, in 1998–99, there were approximately 550,000 fewer stunted children than there would have been if the cumulative distribution function had remained the same as in 1992–93 ($10^9 \times .06 \times .182 \times .05$). In rural areas, the reduction in the number of wasted children amounts to about 1.1 million ($10^9 \times .075 \times .51 \times .03$).

The results for height-for-age (fig. 3) are mixed. In urban areas, child height is improving significantly: the proportion of children with a z-score below -2 or -3 decreases by approximately 3 percentage points, and the distribution for 1998–99 remains below that for 1992–93 for all negative values of z. The p-value of the Kolmogorov-Smirnov test of equality (col. 4 in the central rows in table 3) is .06, and the null of weak first-order stochastic dominance (col. 5) is strongly supported. The intersection-union test (col. 6) rejects the null of no dominance over the range -4 to -1.55. However, in rural areas, our results indicate a striking lack of improvement. The difference between the cumulative distribution functions indicates that there is virtually no change over time in height performances, as confirmed also by the tests in table 3. Hence, while in both sectors measures of short-term nutritional status indicate large improvements, in rural areas, the level of chronic malnutrition appears to have remained overall remarkably stable.

The results described so far exclude Phase I states, that is, Andhra Pradesh,

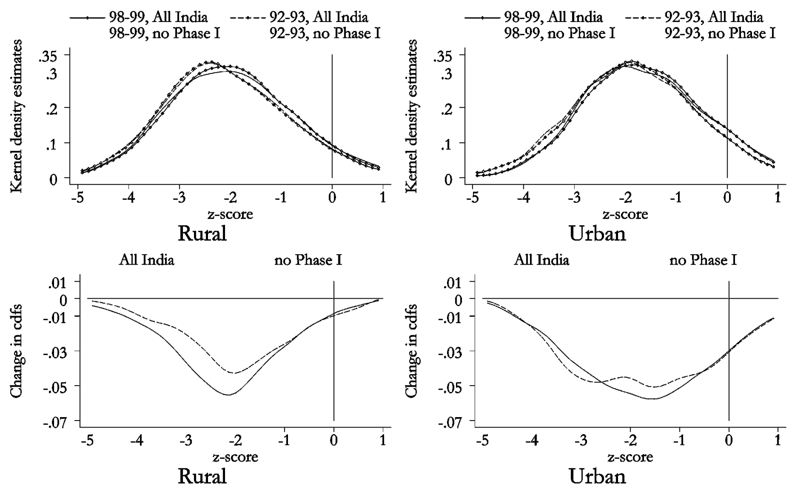

Figure 4. Weight-for-age, both genders, ages 0–3, all India versus all India excluding states with no height in NFHS-I.

Himachal Pradesh, Madhya Pradesh, Tamil Nadu, and West Bengal. Overall, these five states account for approximately 25% of the total Indian population. In figure 4, we compare the changes in distributions for states with nonmissing height with those estimated with the inclusion of Phase I states for the only anthropometric indicator for which such comparison is possible, that is, weight given age. Overall, the two estimated changes are very close in urban areas. In rural areas, the inclusion of Phase I states leads to larger improvements, so that our conjecture is that, in restricting our attention to states for which height was recorded in both rounds, we are underestimating the overall improvements in child nutritional status.[11] For brevity, in the rest of this article, we will abandon weight-for-age as a measure of nutritional status as this indicator is just a combination of height-for-age and weight-for-height.

A. Toward More Gender Inequality?

The results described so far show large improvements in short-term measures of child nutritional status across all of India, as well as sizable improvements in long-term performances in urban areas. In this section, we show that these changes hide large gender differences, especially in rural India. Figure 5 describes sex-specific changes over time in the distribution of weight-for-height and height-for-age. The differences are calculated as before as $F_{98-99}(z) -$

[11] This conjecture is also supported by the observation that the distribution of height-for-age in 1998–99 (not reported here) improves when all states are included. The results are available upon request.

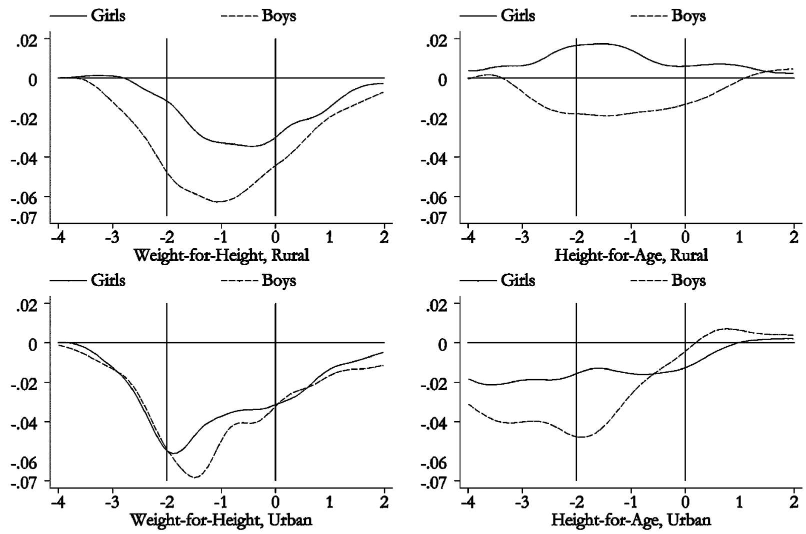

Figure 5. Change in nutritional status over time. Source: authors' calculations from NFHS-I and NFHS-II. All India excluding Andhra Pradesh, West Bengal, Himachal Pradesh, Madhya Pradesh, and Tamil Nadu. Each line is calculated as $F_{II}(z) - F_{I}(z)$.

$F_{92-93}(z)$, so that improvements are represented by negative values. In rural areas, stark gender differences emerge: the change in the distribution of weight-for-height z-scores is about twice as large for boys as for girls, and while boys' height performances show a relatively small improvement (the cumulative distribution function evaluated at -2 drops by approximately .02), for girls we observe an almost specular worsening. These differences are confirmed by the test results reported in table 4. The changes in height are small enough that for both genders the Kolmogorov-Smirnov test of equality does not reject the null of no change at conventional levels (col. 1). The test of stochastic dominance is also not rejected (col. 2), but because the null is of weak stochastic dominance, this result is a simple confirmation of the small changes. The intersection-union tests never reject the null of no first-order stochastic dominance. The tests for boys' weight-for-height strongly support the null of first-order stochastic dominance. The Kolmogorov-Smirnov test for girls gives the same result, while the null of no dominance is not rejected by the intersection-union test, which is more conservative. The bottom two graphs in figure 5 show that changes appear to be much more similar between genders in urban areas. Improvements in boy weight-for-height are large and similar to those observed in rural areas, but girls appear to be doing much better as well; for instance, the proportion with z-scores below -2 decreased by approximately 6 percentage points between 1992–93 and 1998–99 for both boys and girls.

TABLE 4

TESTS OF EQUALITY AND STOCHASTIC DOMINANCE: CHANGES OVER TIME, ALL INDIA, BY GENDER

	Rural			Urban		
	p-Value			*p*-Value		
	Equality (1)	FOSD (2)	No FOSD (3)	Equality (4)	FOSD (5)	No FOSD (6)
Males:						
Height-for-age	.195	.575	NOSD	.015	.72	[−4 − 1.45]
			NOSD			[−4 − 1.45]
			NOSD			[−4 − 3.08]
Weight-for-height	0	1	[−3 1]	0	1	[−3 − .88]
			[−3 1]			[−3 − .96]
			[−3 1]			NOSD
Females:						
Height-for-age	.245	.125	NOSD	.50	.965	NOSD
			NOSD			NOSD
			NOSD			NOSD
Weight-for-height	0	.92	NOSD	.005	1	[−3 − 1.20]
			NOSD			NOSD
			NOSD			NOSD

Sources. Authors' calculations from NFHS-I and NFHS-II, for all India excluding Phase I states.
Note. Columns 1 and 4 report *p*-values of Kolmogorov-Smirnov tests of equality of distributions. In cols. 2 and 5, the null is that the distribution in 1998–99 (weakly) first order stochastically dominates (FOSD) the distribution in 1992–93. Columns 3 and 6 report the results of an intersection-union test of the null of no stochastic dominance (NOSD), using (top to bottom) 10%, 5%, and 1% significance level. See the text for details on the tests. All tests are robust to the presence of intracluster correlation.

However, there is still a gender gap in the change for z-scores between -2 and -1. This is confirmed by the intersection-union tests, which reject the null of no dominance over a larger range (and for smaller significance values) for boys than for girls. Changes in height-for-age are more unequal, as reflected also in the test results reported in table 4: while the cumulative distribution function decreases for girls by about 2 points for z in the interval between -4 and -2, the drop for boys is approximately twice as large.

The changes over time are clearly silent about the gender differences in the cumulative distribution functions in each NFHS round. Because z-scores are normalized using gender-specific growth charts, similar nutritional status for boys and girls relative to the reference growth charts should translate into differences in cumulative distribution functions close to zero. In figure 6, we plot period-specific differences $F_{\text{boys}}(z) - F_{\text{girls}}(z)$, so that we read negative gaps as "boy advantage." In 1992–93 (continuous lines), there is no clear evidence of generalized female disadvantage in nutritional status as evaluated relative to U.S. growth charts.[12] In fact, in both sectors and for almost all values of

[12] Borooah (2005) finds an analogous result with respect to height-for-age, using data collected

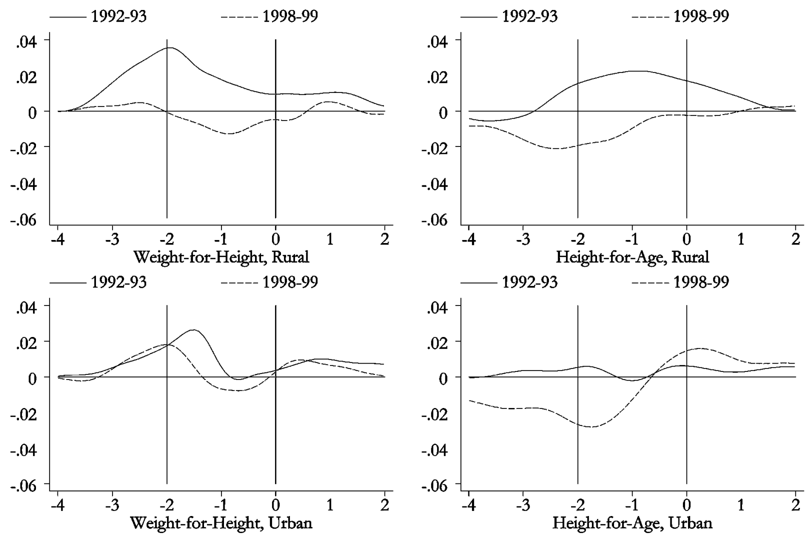

Figure 6. Gender differences in nutritional status. Source: authors' calculations from NFHS-I and NFHS-II. All India excluding Andhra Pradesh, West Bengal, Himachal Pradesh, Madhya Pradesh, and Tamil Nadu. Each line is calculated as $F_{\mathrm{boys}}(z) - F_{\mathrm{girls}}(z)$.

z, growth performances appear to be relatively better for girls. For example, in rural areas, the proportion of girls whose weight-for-height z-score is below -2 is about 3.5 percentage points lower than for boys. In urban areas, differences are generally very small, especially for height-for-age. The curves calculated from NFHS-II (dashed lines) show instead 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 toward male relative advantage in nutritional status. In NFHS-II, the proportion of girls whose weight-for-age z-score is below -2 becomes approximately identical to that for boys. The difference in the proportion of stunted children preserves the same magnitude as in NFHS-I, but the sign is reversed. In urban areas, we observe a small change toward boy advantage over part of the range for weight-for-height and a clear movement toward negative values over much of the range for height.

To analyze whether the change in the gender difference in distributions is not only large in magnitude but also statistically significant, in figure 7 we plot sector-specific "differences-in-differences" of cumulative distribution functions for all India, calculated as

$$\mathrm{DID}(z) = [F^{\mathrm{II}}_{\mathrm{boys}}(z) - F^{\mathrm{II}}_{\mathrm{girls}}(z)] - [F^{\mathrm{I}}_{\mathrm{boys}}(z) - F^{\mathrm{I}}_{\mathrm{girls}}(z)], \quad (1)$$

from the rural areas of the larger Indian states in 1993–94 by the National Council of Applied Economic Research.

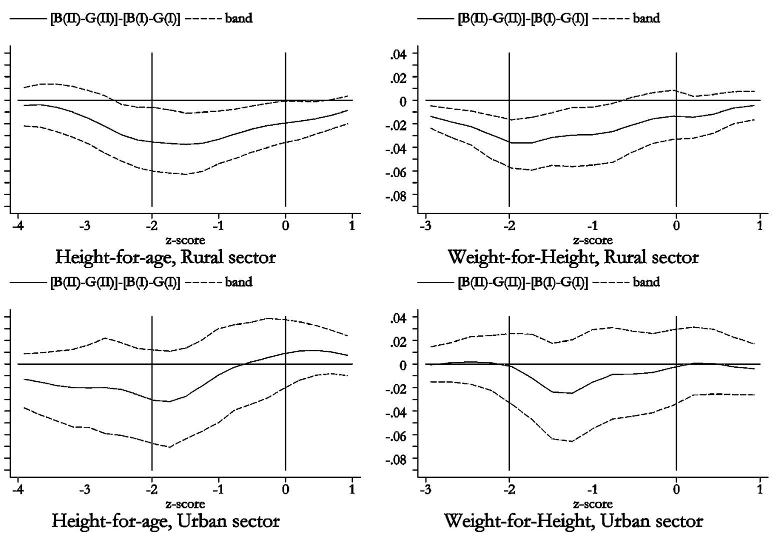

Figure 7. Change over time of gender differences in nutritional status. Source: authors' calculations from NFHS-I and NFHS-II. All India excluding Andhra Pradesh, West Bengal, Himachal Pradesh, Madhya Pradesh, and Tamil Nadu. Each continuous line represents the change over time of the pointwise gender difference in distributions. The dashed lines represent 95% confidence bands (see the text for details).

where the superscript 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 the 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 reestimate 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 includes all observations for a selected cluster, this procedure takes into account both intracluster and intrahousehold correlation, so that the confidence intervals should have correct coverage rates. The confidence bands in figure 7 show that, in rural areas, and especially for weight-for-height, the increase in boy advantage is large, and for most of the relevant range the upper band lies below zero, indicating that over this range the null hypothesis of no change in the gender gap would be rejected at the 5% significance level. In urban areas, the difference-in-differences below zero are also negative, but in this case the bands include zero throughout the whole range.

B. Geographical Differences

In this section, we study the possible existence of geographical patterns in the gender-specific changes in nutritional status. Even if the geographical pattern in the extent of gender inequality is related to social and cultural factors, we do not necessarily expect these factors to affect child outcomes in a time-invariant way. Some of these factors may themselves change, for instance, because of increased female schooling or labor force participation, but (as we have described in the introduction) the way these factors affect child outcomes may also change as a consequence of shifts in economic constraints. Indeed, past research has documented how the extent of gender inequality (as expressed, e.g., in the female-male ratio) has been changing differently in different Indian states (see, e.g., Drèze and Sen 1995, 2002). Ideally, it would be interesting to conduct a separate analysis for each state. However, in order to preserve a relatively large number of observations in each area, we separate India into three broadly defined regions—North, East, and South—following the geo-cultural classification proposed by Sopher (1980). As in the previous section, for comparability reasons, we only include children of up to 3 years of age, 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; Kerala, Karnataka, and Maharashtra represent the South. Table 5 reports the proportion of stunted and wasted children for each sector, gender, and NFHS round, together with the corresponding standard errors and the number of observations used in the calculations. Figures 8–10 display the differences-in-differences estimated as described in Section IV.A.[13] Several striking differences are apparent, both across different regions and between rural and urban areas within the same region.

The results for North India (rows B and F in table 5) are relatively similar to those for the whole country. This is perhaps not surprising as North India accounts for approximately half of all observations (see table 2). In 1992–93, approximately half of children ages 0–3 living in rural areas are stunted, while the proportion is about 5 percentage points lower in urban areas. The extent of wasting is much lower, and wasting affects less than 20% of children in each sector. Notice that, overall, the nutritional status of girls appears to be better than that of boys, at least relative to the growth charts we use as reference. However, in 1998–99, we observe an overturn of this relation, with the exception of weight-for-height in urban areas. The prevalence of stunting in urban areas remains virtually unchanged for girls, while it decreases from

[13] For reasons of space, we omit the graphs for the changes over time in the distributions and for the gender differences in distributions, as well as the tests of stochastic dominance. These additional results are available upon request.

TABLE 5

PROPORTION OF CHILDREN AGES 0–3 WITH A *Z*-SCORE BELOW −2

	1992–93 (NFHS-I)				1998–99 (NFHS-II)			
	Urban		Rural		Urban		Rural	
	Girls (1)	Boys (2)	Girls (3)	Boys (4)	Girls (5)	Boys (6)	Girls (7)	Boys (8)
A. Height-for-age	40.0	40.1	48.6	49.8	38.9	35.6	50.7	48.6
	(1.45)	(1.49)	(.84)	(.84)	(1.41)	(1.23)	(.83)	(.79)
	[2,871]	[3,043]	[7,129]	[7,349]	[2,352]	[2,586]	[6,605]	[7,321]
B. North	43.9	45.1	50.5	50.0	43.3	39.3	55.4	51.9
	(1.95)	(2.03)	(1.13)	(1.12)	(2.02)	(1.65)	(1.08)	(1.08)
	[1,465]	[1,648]	[3,695]	[3,808]	[1,118]	[1,282]	[2,975]	[3,349]
C. East	47.7	41.6	50.0	57.6	42.1	38.7	52.6	51.7
	(4.49)	(4.02)	(1.80)	(1.72)	(3.76)	(4.17)	(1.36)	(1.30)
.	[493]	[478]	[1,471]	[1,499]	[263]	[255]	[1,625]	[1,832]
D. South	32.6	32.3	42.8	39.8	33.0	29.8	40.0	38.3
	(2.09)	(2.59)	(1.76)	(1.64)	(2.18)	(2.09)	(2.22)	(1.98)
	[563]	[526]	[1,124]	[1,226]	[652]	[679]	[857]	[943]
E. Weight-for-height	16.9	18.6	17.7	21.7	11.5	13.6	16.9	16.6
	(1.22)	(.95)	(.64)	(.71)	(.90)	(.88)	(.59)	(.60)
	[2,883]	[3,053]	[7,155]	[7,376]	[2,370]	[2,601]	[6,652]	[7,355]
F. North	16.4	18.1	16.6	19.7	8.2	10.6	12.4	11.8
	(1.50)	(1.32)	(.80)	(.92)	(.95)	(1.12)	(.73)	(.65)
	[1,472]	[1,652]	[3,704]	[3,821]	[1,131]	[1,295]	[3,002]	[3,370]
G. East	11.8	19.5	18.7	25.7	15.9	20.7	20.9	21.6
	(2.42)	(2.66)	(1.41)	(1.62)	(2.39)	(3.15)	(1.10)	(1.11)
	[493]	[479]	[1,476]	[1,501]	[264]	[256]	[1,629]	[1,833]
H. South	19.7	19.5	19.7	21.9	14.8	16.3	21.9	21.0
	(2.50)	(1.69)	(1.55)	(1.27)	(1.71)	(1.46)	(1.46)	(1.63)
	[568]	[530]	[1,133]	[1,231]	[656]	[679]	[868]	[950]

Sources. Authors' calculations from NFHS-I and NFHS-II.
Note. Robust standard errors are reported in parentheses. The number of observations used in the calculation is reported in brackets. Data from Phase I states are excluded.

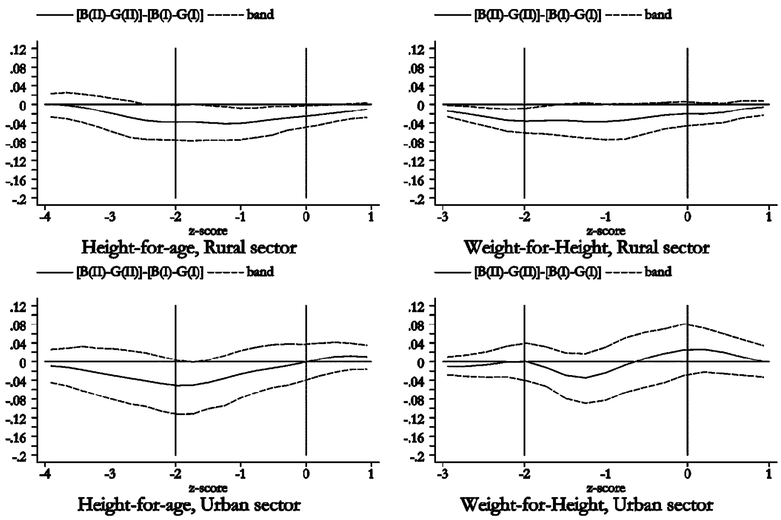

Figure 8. Change over time of gender differences in nutritional status. North India (Gujarat, Haryana, Jammu, Punjab, Rajasthan, Uttar Pradesh, and New Delhi), excluding Phase I states. Each continuous line represents the change over time of the pointwise gender difference in distributions. The dashed lines represent 95% confidence bands (see the text for details).

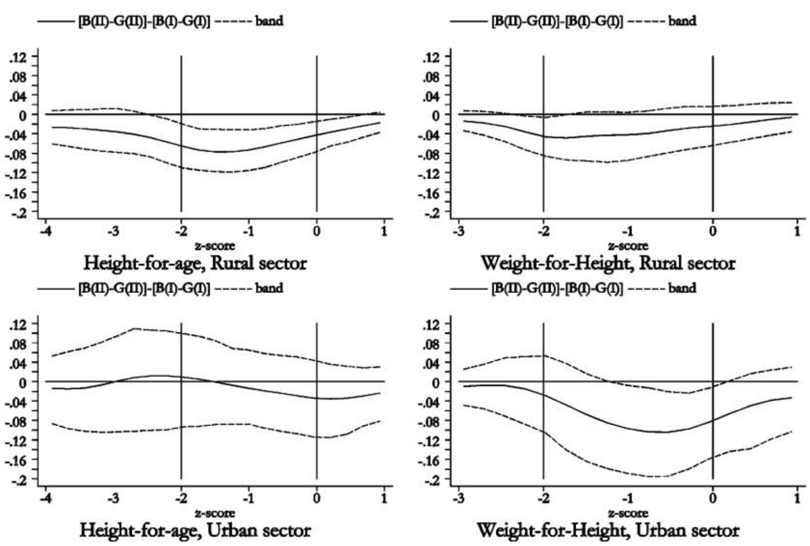

Figure 9. Change over time of gender differences in nutritional status. East India (Assam, Bihar, and Orissa), excluding Phase I states. Each continuous line represents the change over time of the pointwise gender difference in distributions. The dashed lines represent 95% confidence bands (see the text for details).

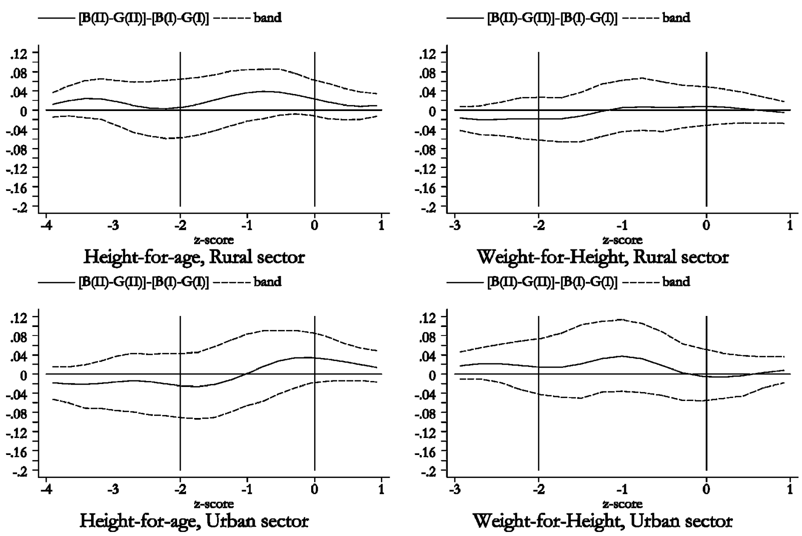

Figure 10. Change over time of gender differences in nutritional status. South India (Kerala, Karnataka, and Maharashtra), excluding Phase I states. Each continuous line represents the change over time of the pointwise gender difference in distributions. The dashed lines represent 95% confidence bands (see the text for details).

45.1% to 39.3% for boys. In rural areas, there is instead an increase in stunting, and the increase for girls (from 50.5% to 55.4%) is more than twice as large as that for boys (from 50% to 51.9%). Looking at wasting (row F), we find instead very large improvements in both sectors, but especially in urban areas, where the proportion below −2 drops by 50% for girls (from 16.4% to 8.2%) and decreases from 18.1% to 10.6% for boys. However, in rural areas, the drop is clearly larger for boys. When we examine changes in the gender differences of the whole distributions (fig. 8), we find a clear movement toward male advantage in the rural sector for both stunting and wasting (even though it is not statistically significant over most of the range). In urban areas the graphs remain below zero for almost all negative z-scores, but the differences are estimated imprecisely, so that the confidence bands always include zero.

In the eastern region (rows C and G in table 5), we observe large improvements in height for both genders in urban areas but only for boys in the rural sector, where the extent of stunting slightly increases among girls. In both NFHS rounds, the extent of stunting is of similar magnitude as for northern states. In urban areas, the proportion of boys who are wasted is slightly higher in 1998–99 (20.7%) than in 1992–93 (19.5%), but the proportion of girls increases considerably from 11.8% to 15.9%. In the rural sector, we find instead an increase in wasting among girls (from 18.7% to 20.9%) and a sizable decrease for boys (from 25.7% to 21.6%). Looking at the changes in the whole distributions, the differences-in-differences in figure 9 show stable gender differences in stunting in urban areas and an important movement toward male advantage in all other cases, where zero remains outside the 95% confidence bands over large sections of the range of z-scores.

The picture for the South, which here includes Kerala, Karnataka and Maharashtra, is completely different. First, note that, while the extent of wasting is roughly comparable to that of other regions, the extent of chronic malnutrition is clearly much lower, so that the proportion of stunted children of either sex is always approximately 30% smaller than in the North or the East. Note also that the degree of stunting remains relatively stable over time. The level of wasting shows instead large improvements in urban areas (where it drops from 19.7% to 14.8% for girls and from 19.5% to 16.3% for boys) and little change in the countryside. The picture emerging from the differences-in-differences (fig. 10) is very different from that in the other region, and in no case do we see a clear movement toward male advantage. The curves always remain close to zero, and over large ranges we even observe positive figures, which indicate a relative gain for girls. However, zero never lies outside the 95% pointwise confidence bands, so that the null of no change in gender differences in the distributions cannot be rejected.

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 benefited much more than girls from a period of rapid economic growth. Only in southern regions and in urban areas elsewhere 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 toward a situation of more pronounced gender inequality in child nutritional status.

V. A First Attempt at Explaining the Changes

In the previous pages, we have shown that the gender differences in the changes in the distribution of z-scores have been markedly different between sectors and across different geographical regions. In this section, we analyze several potential explanations for these trends. First, we consider the role that rural-to-urban migration may have had in shaping the differences between the two different sectors. Then we consider the possible role of changes in boy versus girl mortality in shifting the gender-specific distributions. Finally, we study the relation between changes in child nutritional status and changes over time in a list of economic and demographic variables—defined at the child, household, and community level—that should be strongly associated with child growth performance.

A. Migration

In principle, the observed changes could be explained, at least partly, by migration across different areas rather than by real changes in the nutritional status of children in the relevant age group. For instance, the decline over time of girls' height performances in rural areas and the simultaneous improvement in urban areas (see fig. 5) could be at least partly explained by selective migration of better-off families from rural to urban areas. A similar argument could justify the difference in the improvements in height indicators for boys between rural areas (where small changes are observed) and urban areas (where improvements are more marked).

In this section, we argue that migratory patterns are unlikely to be an important driving force of the observed changes in growth performances. The overall extent of migration in India has been relatively low during recent years, as documented, for instance, in Munshi and Rosenzweig (2005) and Topalova (2005). Topalova (2005), using data from the Indian National Sample Survey, shows that, in 1999–2000, 3.6% of the rural population reported changing either district or sector during the previous 10 years. The proportion is instead higher (13.1%) for urban respondents, but economic considerations are cited as a reason for moving by less than a third of them. Munshi and Rosenzweig (2005), with data from the Rural Economic Development Survey (a representative sample of rural Indian households), estimate that the proportion of men who migrated out of their village of origin was low and actually declined between 1982 and 1999.

However, none of these figures is necessarily very informative about the migration patterns of families with children of ages 0–3. Because most children are born of relatively young parents, who are more likely to migrate, migration rates may be much larger than for the overall population among families where the children in our sample were born. According to the limited information on migration included in the NFHS, in urban areas, 52% of the children of ages 0–3 in 1998-99 were born from mothers who moved to their current residence during the 6 years before the survey while the corresponding figure for rural areas was 41% (table 6). However, the figures are much lower when we look at the proportion who moved and changed sector. The fractions become 25% in urban areas and only 5% in rural areas, where the vast majority of moves are likely to be associated with marriage exogamy. Because the NFHS does not include any information on the state of previous residence, we cannot estimate the fraction of children born of mothers who moved from a different

TABLE 6
MIGRATION

	Gender		
	Male	Female	Total
Urban:			
% moved in 6 years before survey	.52	.53	.52
% moved and changed sector in 6 years before survey	.24	.25	.25
Rural:			
% moved in 6 years before survey	.41	.41	.41
% moved and changed sector in 6 years before survey	.05	.05	.05

Source. NFHS-II (1998–99), all India.
Note. Figures refer to the proportion of mothers of children of ages 0–3 who moved to their current residence during the 6 years before the survey.

state. Note also that there is virtually no difference between the proportion of boys and girls born from mothers who recently changed residence, suggesting that the gender differences in the changes in growth performance are unlikely to be associated with selective migration of parents.

If richer families (who are able to raise taller and better nourished children) moved disproportionately to urban areas and if no real overall improvement in nutritional status took place, we would expect the distributions of z-scores in rural areas to show a decline in nutritional status. However, we observe a decline in growth performances only for girls' height-for-age and not for any other indicators. Also, such form of selective migration would lead us to expect the improvement in the distributions of z-scores for urban children whose mothers did not move between waves to be smaller than the overall change for the urban sector (because, in this scenario, the overall change would also include the inflow of well-fed children). In figure 11, we compare the overall changes over time in the urban gender-specific distribution of z-scores (plotted as continuous lines) with those estimated including only children born from mothers who did not migrate from the rural sector between the two NFHS rounds. There is no evidence that nonmigrants have improved less. If anything, rural-to-urban migration somewhat reduces the overall gain for the urban sectors (especially for girls), suggesting that families that move from rural to urban areas are actually poorer than others who have been living in cities for a long period of time.

B. Changes in Mortality

In this section, we explore the possibility that changes in child mortality may be partly responsible for the gender differences in changes in growth perfor-

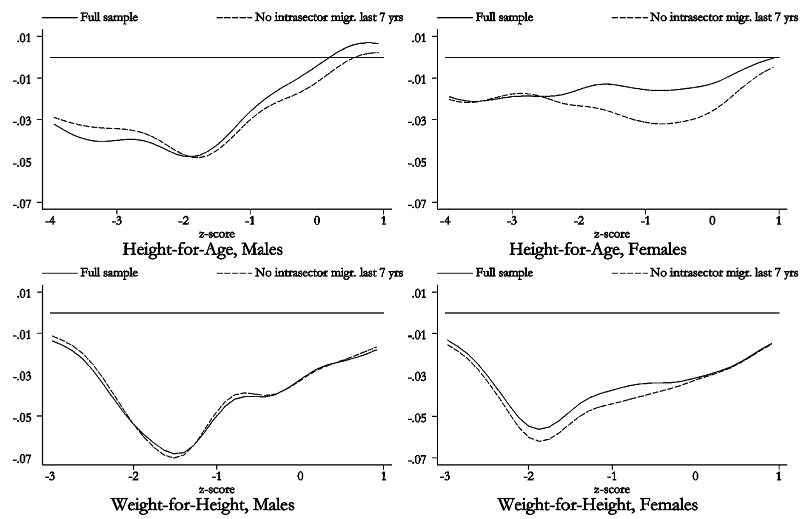

Figure 11. All India, urban, excluding states with no height in NFHS-I

mance that we have documented in Section IV. Our analysis suggests that this possibility is not plausible.

Several studies have documented the existence of a male advantage in the survival probability of young children, a phenomenon that has been especially pronounced in North India. At the same time, we have shown that, in 1992–93, in rural areas of North India and East India, girls appeared to have z-scores no worse (or even better) than boys. In principle, one could reconcile these two results hypothesizing that, in very poor households, resources allocated to girls are not sufficient to guarantee their survival if their nutritional status is below a certain threshold, while son preference is such that boys are taken care of to the extent that they are likely to survive even with very low z-scores. Then the impact of poverty reduction on the distribution of z-scores for girls may have been reduced if the more recent distribution includes girls with very low z-scores who would not have survived in the earlier period. For this argument to be a leading cause of the observed gender differences in changes in nutritional status, we should expect large improvements in girl survival rates and larger improvements for girls than for boys.

Because the NFHS includes a complete birth history, we can estimate survival probabilities for both rounds. For each round, we calculate survival probabilities to age 3, including only data from the 5 years before the survey. This minimizes the likelihood of recall errors and ensures that all births used in the calculations for 1998–99 took place after the conclusion of NFHS-I. We calculate the probability of surviving up to age 3 as $1 - \prod_{i=0}^{3}(1 - q_i)$, where q_0 is neonatal mortality (within the first month of life) and q_i is mortality between $(i - 1)$ and i years of age. For each sector and NFHS round, we estimate separate mortality rates for boys and girls, for all India as well as for each separate geographical region. We also estimate gender differences in mortality rates, and, finally, we estimate both gender-specific changes over time in mortality rates and the changes over time of the gender differences. Because all of these statistics are nonlinear combinations of estimated parameters (the q_i's), we estimate standard errors using 250 bootstrap replications, taking into account the complex survey design. All results are included in table 7.

Mortality rates are clearly very high. In 1992–93, the figures for all India indicate that, in urban areas, for every 1,000 births of a given gender, 67 boys and 71 girls did not survive to the age of 3, while, in rural areas, the figures were 107 and 114, respectively. To put these figures into perspective, the World Health Organization estimated that, in 1992, only 10 children out of 1,000 born in the United States did not survive to the age of 5 (World Health Organization 1995, table 6). The figures for all India show, as expected, higher mortality rates for girls than for boys (see panel C), but such differences are small and not statistically significant at standard confidence levels. However, geographic disaggregation shows that male advantage is larger in the North and the East, while the sign of the differences is reverted in the South. In particular, looking at differences that are statistically significantly different from zero, girls have a 2% higher probability of dying before age 3 in rural areas in the North and a 1.5% lower probability of not surviving in the rural South. Moving to 1998–99, we can see that mortality rates show generalized declines, with the exception of the results for girls in the North and in the rural South and for boys in the rural North. However, the few estimated increases in mortality are very small, and they are not statistically significant. Improvements in mortality rates appear particularly discouraging in North India, where no significant change is observed. Eastern states experienced the largest improvements (especially for girls), with reduction in mortality ranging from −1.4% (for boys in urban areas) to −4% (for girls in the same areas). In the South, mortality rates decreased by slightly more than 1 percentage point in both sectors for boys and in urban areas for girls. Looking at the figures in columns 5 and 6 of panel C, these results do not lend much support to the hypothesis that changes in mortality rates play an important role in explaining the observed gender differences in changes in nutritional status observed in North India and East India (especially in rural areas). In fact, in

TABLE 7
RATES OF CHILD MORTALITY (AGES 0–3) FOR CHILDREN BORN IN THE 5 YEARS PRIOR TO THE SURVEY

	1992–93 (NFHS-I)		1998–99 (NFHS-II)		Change	
	Urban (1)	Rural (2)	Urban (3)	Rural (4)	Urban (5)	Rural (6)
A. Male:						
India	.067	.107	.061	.093	−.0066	−.0149
	(.0039)	(.0027)	(.0039)	(.0024)	[.2289]	[.0001]*
North	.072	.114	.074	.106	.0011	−.0083
	(.0055)	(.0042)	(.0067)	(.0038)	[.8997]	[.1418]
East	.077	.114	.063	.096	−.0137	−.0186
	(.0111)	(.0050)	(.0104)	(.0046)	[.3578]	[.0118]*
South	.058	.088	.047	.067	−.0108	−.0226
	(.0067)	(.0057)	(.0066)	(.0057)	[.2331]	[.0034]*
B. Female:						
India	.071	.114	.060	.105	−.0114	−.0092
	(.0045)	(.0030)	(.0040)	(.0032)	[.0688]	[.0397]*
North	.076	.133	.075	.133	.0000	.0002
	(.0069)	(.0052)	(.0059)	(.0049)	[.9975]	[.9771]
East	.094	.118	.055	.087	−.039	−.032
	(.0113)	(.0067)	(.0090)	(.0042)	[.0045]*	[.0000]*
South	.056	.07	.043	.077	−.0122	.002
	(.0078)	(.0052)	(.0071)	(.0057)	[.2421]	[.7689]
C. Differential (m(girls) − m(boys)):						
India	.0039	.0068	−.0017	.0124	−.0056	.0056
	(.0057)	(.0041)	(.0049)	(.0038)*	[.456]	[.316]
North	.0041	.0196	.0014	.027	−.002	.008
	(.0083)	(.0063)*	(.0072)	(.0060)*	[.806]	[.352]
East	.0177	.0040	−.0073	−.0095	−.025	−.013
	(.0161)	(.0080)	(.0105)	(.0059)	[.193]	[.174]
South	−.0017	−.0141	−.0033	.0102	−.0016	.0243
	(.0091)	(.0068)*	(.0094)	(.0079)	[.903]	[.020]*

Sources. Authors' calculations from NFHS-I and NFHS-II.
Note. Standard errors are in parentheses in cols. 1-4; *p*-values for the test of equality are in brackets in cols. 5 and 6. All standard errors and tests take into account clustering and stratification (at the state level). All standard errors are calculated using 250 bootstrap replications. For details on the calculation of the mortality rates, see Sec. V.B.
* In cols. 1–4 of panel C, this indicates that the null hypothesis of no difference in mortality rates between boys and girls is rejected at the 5% significance level. In cols. 5 and 6 of all panels, this indicates that the null hypothesis of no change over time in the corresponding statistic is rejected at the 5% significance level.

northern states (which account for about half of the children in the sample), there is virtually no change in the probability of survival up to age 3. In eastern states, survival probabilities do increase more for girls than for boys, but the difference is too small to explain the large difference-in-differences in cumulative distribution functions observed in rural areas (see the bottom panel of fig. 9); for instance, while the fraction of girls surviving to age 3 increased by 1.4 percentage points more than the corresponding fraction for boys, the difference in the proportion of boys versus girls with a height-for-age z-score below -2 decreased by approximately 5 percentage points.

C. Looking for Factors Driving the Changes in Nutritional Status

In the previous sections, we have examined if changes in child mortality or migration patterns from rural to urban areas can mechanically explain at least part of the trends described in Section IV, and we have argued that this does not appear to be the case. In this section, we attempt to evaluate how much of the changes can be explained, first, by changes over time in the distribution of economic and demographic factors that are likely to be important predictors of child nutritional status and, second, by changes in the "returns" of these factors on child nutrition. For the first 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). Namely, given the cumulative distribution function $F(z)$ of an anthropometric index z and letting $\boldsymbol{x}$ denote a vector of predictors, we estimate a counterfactual distribution of z for 1998–99 using the conditional distribution $F(z \mid \boldsymbol{x})$ in 1992–93 and $F(\boldsymbol{x})$ in 1998–99. On the one hand, this semiparametric approach has the advantage of analyzing changes in the whole distribution, but, on the other hand, its nonparametric nature is not easily adapted to studying changes in the conditional relation between child nutrition and its predictors. Hence, for this second purpose, in Section V.D, we make use of conventional Oaxaca decompositions, and we shift focus from the whole distribution to the more limited analysis of the probability of stunting. In this section, we only discuss the results for height-for-age, as the explanatory factors in our analysis are likely to be less useful in explaining a short-term measure of nutritional status such as weight-for-height.[14]

The list of predictors of child nutritional status that we use includes a series of child-, household-, and village-specific variables. We exclude such variables as housing characteristics, labor supply, and asset ownership (the NFHS does not include information on expenditure or income). Even though these latter variables are likely to be good predictors of child nutritional status, they are certainly endogenous as they are largely determined jointly with expenditures for child nutrition and health care. We include a polynomial in age to capture the very strong association between age and z-scores (see fig. 1). This allows us to evaluate if part of the change in the distribution of z-scores is simply due to a change in the age distribution of children. Father's and (especially) mother's education have been widely documented as important determinants of child health (e.g., see Drèze and Sen 2002, chap. 7 and references therein). Mother's schooling is categorized with dummies for mother illiterate (omitted), mother literate below middle school, completed middle school, and high school

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

and above. For father's education, we use dummies for no schooling or at least 12 years of completed schooling (which corresponds to the attainment of a secondary school degree). Due to the limited scope of the land market in India, we can also use a dummy for land ownership as an exogenous predictor.[15] The household demographic structure is taken into account by including household size and a dummy for high birth order set to be equal to one for children with more than three older siblings. Among the household characteristics, we also include religion (dummies for Muslim and "other religions," with the Hindu category omitted) and a binary variable equal to one when the household head is a woman. Information on caste is included in both surveys, but we chose not to use it because definitions are not consistent between the two questionnaires. Finally, for children living in rural areas, we use a set of indicators for community characteristics. These village-level variables are only available for the rural sample, for which both surveys also include a "village questionnaire." In our choice of community characteristics, we are limited by a number of noncomparability issues due to differences in the variables included in the village questionnaire of the two surveys. Ultimately, we are left with the use of a list of binary variables equal to one if the following are present in the village: electrification, fair price shop, no drainage, *Anganwadi, Mahila Mandal,* pharmacy, health subcenter, and primary health center.[16] The inclusion of measures of health facilities may be important in explaining changes in child nutritional status. For instance, Deolalikar (2005) stresses the role that increased government spending on health and nutrition program should have in reaching targets of reduced malnutrition and child mortality in India.[17] It should be noted that the use of community variables has the drawback of leading to a reduction in the rural sample size as in both surveys several observations are missing and that, in 1992–93, we lose some villages for which the village questionnaire cannot be matched to the individual data. Overall,

[15] We do not use the information on land cultivated and irrigated included in the two NFHS waves as these variables are recorded differently in the two surveys. Moreover, while land ownership can often be assumed to be exogenous in India, it is less clear that this assumption can be used for land cultivated and irrigated.

[16] Fair price shops are special retail shops where subsidized staples offered through the Indian Public Distribution System can be purchased by eligible households; *Anganwadis* are child-care centers that operate as the focal point for the delivery of services at the community level to children below 6 years of age, pregnant and nursing mothers, and adolescent girls; *Mahila Mandals* (women's clubs) are village women associations that also have the purpose of sharing health knowledge among members; primary health centers are local health centers that also supervise the operation of more subcenters, which serve a smaller number of families.

[17] Note, however, that recent research has carefully documented how the existence of health structures is far from sufficient to guarantee the provision of effective health services. See, e.g., Duflo, Banerjee, and Deaton (2004) and Das and Hammer (2005).

missing data lead to the loss of approximately 16% of the rural sample in NFHS-I and 7% of the rural sample in NFHS-II.

We turn now to the description of the estimation of the counterfactual distributions. We describe the estimation for the case where the covariates $\boldsymbol{x}$ are continuous, but, with a change of notation, the argument can be straightforwardly adapted to the case where some of the covariates are discrete. Formally, let $f(\bar{z} \mid t)$ be the true density of the anthropometric index z evaluated at $\bar{z}$ in wave t, where $t = \mathrm{I}, \mathrm{II}$. The density can be rewritten as

$$f(\bar{z} \mid t) = \int f(\bar{z} \mid \boldsymbol{x}, t) f(\boldsymbol{x} \mid t) d\boldsymbol{x}, \tag{2}$$

where $f(\boldsymbol{x} \mid t)$ is the density of the covariates $\boldsymbol{x}$ in wave t. For notational convenience, let us write $f(\bar{z} \mid t) \equiv f(\bar{z} \mid t_{\boldsymbol{x}} = t, t_{z|\boldsymbol{x}} = t)$, where $t_{z|\boldsymbol{x}}$ indicates the wave that identifies the conditional distribution of z given $\boldsymbol{x}$, and $t_{\boldsymbol{x}}$ indicates the wave that identifies the marginal distribution of $\boldsymbol{x}$. Clearly, the two waves coincide in the actual density. We are interested in studying how much of the changes in the distribution of z-scores can be explained by changes in the distribution of the covariates, keeping the distribution of z conditional on $\boldsymbol{x}$ constant. In other words, we want to estimate the counterfactual density $f(\bar{z} \mid t_{\boldsymbol{x}} = \mathrm{II}, t_{z|\boldsymbol{x}} = \mathrm{I})$. A straightforward way to estimate this object follows after noting that it can be usefully rewritten as follows:

$$f(\bar{z} \mid t_{\boldsymbol{x}} = \mathrm{II}, t_{z|\boldsymbol{x}} = \mathrm{I}) = f(\bar{z} \mid t = \mathrm{I}) E[R(\boldsymbol{x}) \mid \bar{z}, t = \mathrm{I}], \tag{3}$$

where

$$R(\boldsymbol{x}) = \frac{P(t_{\boldsymbol{x}} = \mathrm{II} \mid \boldsymbol{x}) P(t_{\boldsymbol{x}} = \mathrm{I})}{P(t_{\boldsymbol{x}} = \mathrm{I} \mid \boldsymbol{x}) P(t_{\boldsymbol{x}} = \mathrm{II})}. \tag{4}$$

The proof follows from a straightforward application of the properties of probabilities. The function $R(\boldsymbol{x})$ is a "reweighting function" that maps the conditional density from Wave I into the counterfactual density $f(\bar{z} \mid t_{x} = \mathrm{II}, t_{z|\boldsymbol{x}} = \mathrm{I})$ by increasing (decreasing) the contribution to this counterfactual marginal density of the conditional density $f(\bar{z} \mid \boldsymbol{x}, t_{z|\boldsymbol{x}} = \mathrm{I})$ for values of $\boldsymbol{x}$ that are relatively common (rare) in Wave II. The different components of the reweighting function are estimated pooling together data from both NFHS waves. Then the unconditional probability $P(t_{\boldsymbol{x}} = \mathrm{I})$ can be simply estimated as the (weighted) fraction of observations that belongs to the first wave, while the conditional probability $P(t_{\boldsymbol{x}} = \mathrm{II} \mid \boldsymbol{x})$ can be interpreted as the probability that an observation with covariates equal to $\boldsymbol{x}$ belongs to the second NFHS wave, and it can be estimated using a binary dependent variable model.

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

$$\hat{f}(\bar{z} \mid t_{\boldsymbol{x}} = \mathrm{II}, t_{z|\boldsymbol{x}} = \mathrm{I}) = \sum_{i \in I} w_i \hat{R}(\boldsymbol{x}_i) \frac{1}{h} K\left(\frac{\bar{z} - z_i}{h}\right), \tag{5}$$

where w_i is the sampling weight for the ith observation (normalized so that $\sum_{i \in I} w_i = 1$), $K(.)$ is a standard kernel, h is the bandwidth, and $i \in \mathrm{I}$ 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.[18]

For the sake of brevity, here we only report the results for all Indian states (excluding as usual Phase I states). We show the resulting predicted changes in figure 12. For both sectors, we present four different lines: the actual change, the change in the distribution predicted by the sole change in the distribution of demographic variables (age, household size, and high birth order), the change predicted by the sole change in parental education, and, finally, the predicted change estimated including all predictors.

A few conclusions emerge. First, in all cases, the change in the age distribution and other demographic variable predict virtually no improvement in child nutritional status, so that the observed changes are not the result of a mere change in the distribution of child age. Second, in urban areas, the changes in height predicted by improvements in parental schooling are not too dissimilar to the actual ones, which are, however, larger for low z-scores. The inclusion of the complete set of predictors leaves the results almost unaffected. In rural areas, changes in parental education, as well as changes in all included variables, predict a small improvement in girl height performances. This contrasts with the small worsening observed instead in the data. The actual small improvement in boy height is very close to the change predicted by the increase in parental education. However, the inclusion of community variables among the predictors, unlike for girls, increases the predicted decline in the cumulative distribution function by approximately 1 percentage point for all negative z-scores. It is interesting that, in both rural and urban areas, our prediction exercise forecasts much larger improvements for boys than for girls, suggesting that at least part of the gender gap in the changes in nu-

[18] In principle, one can estimate directly the counterfactual cumulative distribution functions using a procedure analogous to that just described (see Tarozzi [forthcoming] for details). We chose to estimate the densities first because the resulting graphs are smoother.

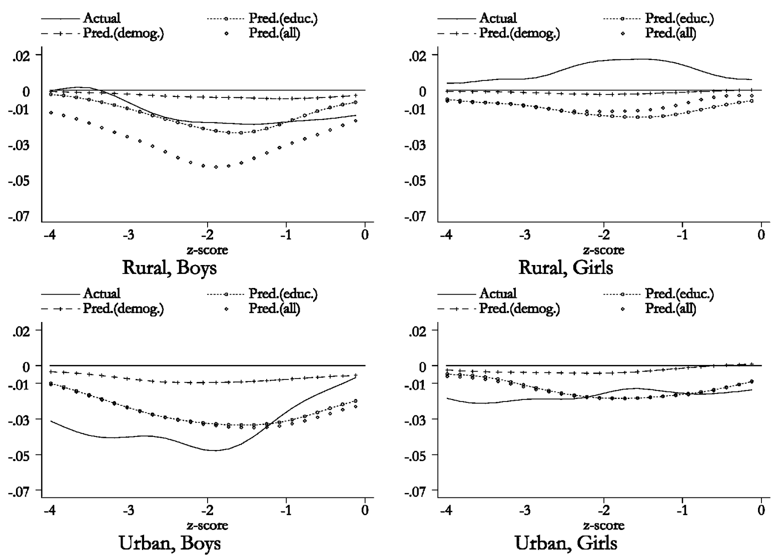

Figure 12. Counterfactual distributions, height-for-age

tritional status over time may be due to an association between growth performance and predictors that is stronger for boys than for girls.

D. Oaxaca Decompositions

On the one hand, part of the discrepancy between predicted and actual changes documented in the previous subsection is certainly due to the relatively short list of predictors included in our analysis. On the other hand, our results suggest that changes in the distribution of z-scores conditional on the predictors are likely to have changed over time. The semiparametric approach used so far has the advantage of analyzing changes in the whole distribution, but its nonparametric nature is not easily adapted to studying changes in the conditional relation between child nutrition and its predictors. Hence, for this second purpose, we make use of conventional Oaxaca decompositions (Oaxaca 1973) applied to the analysis of the probability of stunting, which we estimate with linear probability models. We choose to use a linear probability model even if the dependent variable is binary because a linear model makes the decomposition results easier to interpret. Moreover, the slopes estimated with linear probability models are usually very close to the marginal effects routinely estimated for binary dependent variable models such as logit or probit.[19]

The use of Oaxaca decompositions allows us to examine how the contribution of different factors to boy and girl height is changing over time in different

[19] See Nielsen (1998) for a decomposition technique appropriate for logit models.

geographical areas. In Section IV, we showed that the gender differences in changes over time are particularly striking in rural areas of North India and East India. For this reason, and also to save space, here we pool together the North and the East and restrict our analysis to the rural sector, which accounts for the majority of the population and is where the extent of stunting is larger than in urban areas. Results for the urban sector are available upon request.

Let D^z_{igs} denote a dummy equal to one if the z-score of child i of gender $g, g = m, f$, measured in survey $s, s = \mathrm{I}, \mathrm{II}$ is below -2. Let X'_{igs} denote the vector of determinants of child nutritional status. Then the model for a given gender and wave is

$$D^z_{igs} = X'_{igs}\beta_{gs} + \epsilon_{igs}, \tag{6}$$

so that the change in means over time can be decomposed as

$$\begin{aligned} \bar{D}^z_{gII} - \bar{D}^z_{g\mathrm{I}} &= \bar{X}'_{g\mathrm{II}}\hat{\beta}_{g\mathrm{II}} - \bar{X}'_{g\mathrm{I}}\hat{\beta}_{g\mathrm{I}} \\ &= (\bar{X}_{g\mathrm{II}} - \bar{X}_{g\mathrm{I}})'\hat{\beta}_{g\mathrm{I}} + \bar{X}'_{g\mathrm{II}}(\hat{\beta}_{g\mathrm{II}} - \hat{\beta}_{g\mathrm{I}}). \end{aligned} \tag{7}$$

The first term in (7) can be interpreted as the part of the change in the mean of the dependent variable associated to a change in the means of the regressors, while the second term is the part due to a change in the coefficients.[20]

We report the results of the decompositions in tables 8 and 9. Each table includes separate decompositions for males and females for a given geographical area. For each gender, the first two columns display the estimated coefficients of the linear probability model. For each listed predictor, the third column reports $\Delta_{X_i} \equiv (\bar{X}_{i,g\mathrm{II}} - \bar{X}_{i,g\mathrm{I}})\hat{\beta}_{i,g\mathrm{I}}$, that is, the change in the probability of stunting predicted by the change in the mean value of the predictor, keeping all estimated coefficients equal to their values in 1992–93 and keeping all other regressors' means constant. The figures in the last column represent instead $\Delta_{\beta_i} \equiv \bar{X}_{i,g\mathrm{II}}(\hat{\beta}_{i,g\mathrm{II}} - \hat{\beta}_{i,g\mathrm{I}})$, that is, the change predicted by the shift in the estimated coefficients. Each regression also includes (not shown) a constant and a cubic in age and household size. For each decomposition, we also report the total change predicted by the change in the mean value of the predictors and the total change predicted by the change in coefficients. By construction, these two totals (denoted "Total" in the tables) sum up to the change over time in the proportion of stunted children. Finally, below these total changes, we calculate the changes predicted based on subsets of regressors, namely, the

[20] Note that even if the two components, by construction, have to sum up to the change in the mean dependent variable, their signs may differ, so that the fraction of the change associated to each component is not constrained to be bounded between zero and one.

three maternal education dummies (Mo.Ed. in the tables), the two paternal education dummies (Fa.Ed.), and the variables that measure the availability of health-related amenities in the village (Health). Note that, even though the regressions do not include such variables as income or asset ownership (which are certainly endogenous because determined jointly with child health inputs), one should be very cautious in interpreting the regression results in a causal way. Most of the included regressors are, in fact, likely to be correlated with unobserved heterogeneity in preferences, cultural norms, or other location-specific characteristics that may also have a direct impact on the dependent variable. Similarly, endogenous placement of village amenities such as health structures or fair price shops can further hinder the causal interpretation of the corresponding coefficients. For these reasons, we think that the interest of these results lie more in their descriptive content than in their causal meaning, which is, at best, doubtful.

In table 8, we examine the decompositions for the rural North and East of the change in stunting. The proportion of stunted children decreases from 52.9% to 51.8% for boys, while it increases from 51.2% to 54.6% for girls. The R^2s show that the regressors predict approximately 20% of the variance of the dependent variable. For both boys and girls, maternal schooling is strongly associated with lower levels of stunting, and in both waves, higher schooling leads to larger reductions. Moreover, the magnitude of the coefficients is always larger in 1998–99 than in 1992–93; for instance, in NFHS-I, a boy whose mother has completed middle school (high school or above) has a 7.8% (15.5%) smaller probability of being stunted than if he had an illiterate mother, and the reduction becomes 13.9% (24.5%) in 1998–99. Similar patterns emerge for girls, even if in every single case mother's education is associated with a larger reduction in stunting for boys than for girls. Overall, there is a .8% predicted decline in stunting among boys due to increases in maternal education, to which a further 1.3% decline is added due to the increased magnitude of the coefficients. For girls, the two figures are, respectively, −.4% and −1.2%. Overall, then, maternal education appears to have benefited boys only marginally more than girls. A large contribution to the gender difference in the change in the extent of stunting between the two surveys comes from the health-related community variables. First, the change in their availability predicts an overall 1.8% decline in stunting for boys but only a .1% drop for girls. Second, the change in the corresponding coefficients sums up to a 1.2% increase in stunting among girls, while it predicts an almost 3% decline for boys. Looking at all regressors, the change in their mean value predicts large improvements for boys (−.039) and small gains for girls (−.002).

In the South (see table 9), the fraction of stunted children decreases from

39.8% to 37.8% for boys, while it decreases more for girls, from 43.5% to 39.8%. As in North India and East India, maternal education is strongly negatively associated with stunting, but in the South there is evidence of a strong increase in the "returns" to mother's schooling only for girls whose mothers have at least a middle school diploma. The change in maternal schooling leads to similar predicted improvements between genders (a 2.2% decline in stunting for boys and a 1.8% decline for girls). The change in the corresponding slopes increases stunting for boys by 1.6%, while it decreases it by 1.7 percentage points for girls. Increases in father's education predict approximately identical decreases in stunting for boys and girls, and the sum of the corresponding Δ_{β_i} further reduces stunting by .5 percentage points for boys and by 2.9% for girls. The contribution of community characteristics is somewhat difficult to explain. The change in the level of the regressors predicts a 1.5% decline in stunting for boys and almost a zero decline for girls (a result very similar to what we observed in the North and the East), but the coefficient changes sum up to a 1.5% increase in stunting for girls and a 20% increase for boys, which is almost completely explained by the change in the coefficients for availability in the village of drainage, *Anganwadis*, and *Mahila Mandals*. Notably, while the presence of an *Anganwadi* is associated with a 7% decline in boy stunting in 1992–93 (significant at the 5% level), the coefficient becomes even larger but of opposite sign (but not significant) in 1998–99. Note, however, that this very large figure (20%) is more than compensated for by a sizable increase in the magnitude of the negative coefficient that relates stunting to village electrification. Overall, the decline in stunting predicted by the change in the covariates is very similar between genders (Δ_X is $-.038$ for boys and $-.032$ for girls). Most of the difference (in favor of girls) arises from the changes in the slopes, which almost average out to zero for girls ($\Delta_\beta = .005$), while the slope changes dampen the reduction in stunting for boys ($\Delta_\beta = .018$).

VI. Conclusions

The Indian National Accounts show rapid rates of gross domestic product (GDP) growth during the 1990s. Estimates on the reduction in poverty during this decade are not unanimous, but according to several researchers poverty declined considerably, especially in urban areas. In this article, we use data from two independent cross-sectional surveys (completed in 1992–93 and 1998–99) to evaluate to what extent the growth observed during the nineties has been associated with a reduction in malnutrition among children ages 0–3. We find that measures of short-term nutritional status based on weight given height show large improvements, especially in urban areas. For instance,

TABLE 8

OAXACA DECOMPOSITION, HEIGHT-FOR-AGE, RURAL NORTH-EAST

	$\beta_{92\text{–}93}$	$\beta_{98\text{–}99}$		Δ_X	Δ_β
Male:					
Mother literate below middle school	−.0659 (−2.45)	−.0732 (−3.41)		−.0013	−.0010
Mother completed middle school	−.0777 (−1.90)	−.1387 (−4.85)		−.0023	−.0047
Mother completed at least high school	−.1545 (−4.07)	−.2454 (−8.14)		−.0043	−.0076
Father has no formal education	−.0214 (−1.11)	.0507 (3.13)		.0012	.0241
Father has at least 12 years education	−.0712 (−2.32)	.0055 (.20)		−.0012	.0107
High birth order (above 3)	.0479 (2.88)	−.0014 (−.08)		−.0018	−.0171
Household head is woman	−.0339 (−1.11)	−.0055 (−.17)		−.0002	.0015
Muslim	.0044 (.16)	−.0114 (−.49)		−.0001	−.0022
Other religions	−.1455 (−4.76)	−.0732 (−2.09)		.0019	.0029
Health subcenter in village	−.0308 (−1.42)	−.0083 (−.46)		−.0020	.0074
Primary health center in village	−.0279 (−.80)	.0024 (.09)		−.0011	.0034
Village has no drainage	.0465 (2.29)	−.0078 (−.49)		−.0017	−.0305
Anganwadi in village	−.0413 (−2.02)	−.0109 (−.64)		−.0076	.0163
Mahila Mandal in village	.0312 (1.38)	−.0256 (−1.35)		−.0020	−.0112
Pharmacy in village	.0645 (2.87)	−.0044 (−.21)		−.0040	−.0147
Fair price shop in village	−.0295 (−1.48)	.0025 (.15)		−.0056	.0173
Village is electrified	−.0356 (−1.54)	−.0206 (−1.05)		−.0019	.0109
Household owns land	−.0156 (−.85)	−.0230 (−1.40)		.0008	−.0050
			Total	−.0392	.0281
N	4,378	4,866	Mo.Ed.	−.0079	−.0132
R^2	.1535	.1704	Fa.Ed.	.0000	.0347
$\hat{P}(z < -2)$	.5293	.5183	Health	−.0184	−.0293

Female:					
Mother literate below middle school	−.0329 (−1.35)	−.0647 (−3.02)		−.0004	−.0044
Mother completed middle school	−.0425 (−1.07)	−.0919 (−2.64)		−.0005	−.0028
Mother completed at least high school	−.1407 (−3.58)	−.2047 (−6.08)		−.0030	−.0049
Father has no formal education	.0052 (.29)	.0115 (.65)		−.0004	.0021
Father has at least 12 years education	−.0288 (−1.05)	−.0445 (−1.64)		−.0004	−.0020
High birth order (above 3)	.0366 (1.98)	.0102 (.59)		−.0012	−.0092
Household head is woman	.0019 (.05)	.0133 (.38)		.0000	.0006
Muslim	.0376 (1.57)	−.0241 (−.97)		−.0007	−.0079
Other religions	−.1100 (−3.75)	−.0799 (−1.92)		.0016	.0011
Health subcenter in village	−.0302 (−1.35)	−.0087 (−.47)		−.0026	.0073
Primary health center in village	−.0166 (−.51)	−.0009 (−.03)		−.0006	.0018
Village has no drainage	−.0107 (−.52)	−.0004 (−.02)		.0005	.0057
Anganwadi in village	.0044 (.22)	.0145 (.82)		.0008	.0054
Mahila Mandal in village	−.0393 (−1.83)	−.0378 (−1.76)		.0027	.0003
Pharmacy in village	.0428 (1.95)	.0025 (.11)		−.0022	−.0089
Fair price shop in village	.0277 (1.44)	−.0251 (−1.43)		.0058	−.0291
Village is electrified	−.0476 (−2.14)	−.0259 (−1.29)		−.0028	.0159
Household owns land	−.0268 (−1.43)	−.0184 (−1.06)		.0007	.0057
			Total	−.0019	.0359
N	4,280	4,313	Mo.Ed.	−.0040	−.0121
R^2	.183	.2182	Fa.Ed.	−.0008	.0001
$\hat{P}(z<-2)$	.5116	.5457	Health	−.0013	.0116

Sources. Authors' calculations from NFHS-I and NFHS-II.

Note. All India except Phase I states. Clustered *t*-ratios are in parentheses. Each regression also includes (not shown) a constant and a cubic in age and household size. The first two columns display the coeffficients estimated in the linear probability model. The third column reports the change in the probability of stunting predicted by the change in the mean value of the row-specific predictor. The figures in the last colum indicate the change predicted by the shift in the estimated coefficients. The sum of the changes are reported as "Total," and below these we calculate the changes predicted by the three maternal education dummies (Mo.Ed.), the two paternal education dummies (Fa.Ed.), and the variables that measure the availability of health-related amenities in the village (Health). See the text for details.

TABLE 9

OAXACA DECOMPOSITION, HEIGHT-FOR-AGE, RURAL SOUTH

	β_{92-93}	β_{98-99}		Δ_X	Δ_β
Male:					
Mother literate below middle school	−.0710 (−1.78)	−.0896 (−1.67)		.0060	−.0037
Mother completed middle school	−.1715 (−3.44)	−.1251 (−2.31)		−.0080	.0067
Mother completed at least high school	−.1709 (−2.94)	−.1136 (−1.97)		−.0203	.0128
Father has no formal education	.0759 (2.23)	.1039 (2.47)		−.0035	.0066
Father has at least 12 years education	−.0267 (−.41)	−.0970 (−2.07)		−.0025	−.0112
High birth order (above 3)	.0109 (.32)	.1315 (2.55)		−.0006	.0215
Household head is woman	−.0242 (−.64)	.0509 (.88)		.0008	.0059
Muslim	.0186 (.46)	−.1168 (−2.53)		.0000	−.0205
Other religions	.1163 (2.15)	−.1250 (−1.84)		−.0024	−.0129
Health subcenter in village	−.0500 (−1.29)	−.0156 (−.35)		.0028	.0167
Primary health center in village	−.0597 (−1.20)	−.0093 (−.19)		−.0056	.0161
Village has no drainage	−.0706 (−2.42)	.0155 (.40)		.0001	.0327
Anganwadi in village	−.0707 (−2.18)	.0712 (1.25)		−.0142	.1275
Mahila Mandal in village	.0335 (1.05)	.0715 (1.91)		.0012	.0255
Pharmacy in village	.0395 (.80)	−.0039 (−.09)		.0000	−.0152
Fair price shop in village	−.0662 (−1.89)	−.0529 (−1.25)		.0037	.0097
Village is electrified	.1625 (.86)	−.0764 (−.68)		−.0045	−.2290
Household owns land	.0277 (.84)	−.0260 (−.74)		−.0014	−.0330
			Total	−.0375	.0175
N	1,148	872	Mo.Ed.	−.0222	.0159
R^2	.2087	.1806	Fa.Ed.	−.0060	−.0047
$\hat{P}(z<-2)$	.3975	.3775	Health	−.0156	.2034

Female:					
Mother literate below middle school	−.1302 (−3.18)	−.0925 (−1.76)		−.0023	.0093
Mother completed middle school	−.0680 (−1.34)	−.1780 (−2.50)		−.0019	−.0152
Mother completed at least high school	−.1463 (−2.72)	−.2031 (−3.07)		−.0142	−.0115
Father has no formal education	.0418 (1.12)	−.0501 (−.95)		−.0021	−.0248
Father has at least 12 years education	−.0543 (−.97)	−.0798 (−1.36)		−.0037	−.0039
High birth order (above 3)	.0604 (1.54)	.0539 (1.07)		−.0028	−.0011
Household head is woman	−.0698 (−1.51)	−.0644 (−1.11)		.0012	.0004
Muslim	−.0163 (−.38)	−.0425 (−.83)		−.0003	−.0039
Other religions	−.0124 (−.21)	−.1529 (−2.50)		.0003	−.0083
Health subcenter in village	−.0791 (−2.20)	.0049 (.11)		.0052	.0397
Primary health center in village	−.0444 (−1.11)	−.0256 (−.41)		−.0042	.0058
Village has no drainage	.0300 (1.00)	.0000 (.00)		.0012	−.0127
Anganwadi in village	−.0368 (−1.02)	−.0022 (−.03)		−.0062	.0310
Mahila Mandal in village	.0463 (1.37)	−.0007 (−.02)		.0022	−.0317
Pharmacy in village	−.0033 (−.07)	−.0501 −.91		−.0001	−.0173
Fair price shop in village	.0375 (.98)	−.0068 (−.14)		−.0015	−.0333
Village is electrified	−.0680 (−.56)	−.1433 (−1.02)		.0015	−.0728
Household owns land	.0158 (.46)	−.0251 (−.63)		−.0006	−.0251
			Total	−.0321	−.0045
N	1,055	785	Mo.Ed.	−.0184	−.0174
R^2	.2562	.1737	Fa.Ed.	−.0058	−.0286
$\hat{P}(z<-2)$	.4346	.3980	Health	−.0019	.0147

Sources. Authors' calculations from NFHS-I and NFHS-II.

Note. All India except Phase I states. Clustered *t*-ratios are in parentheses. Each regression also includes (not shown) a constant and a cubic in age and household size. The first two columns display the coeffficients estimated in the linear probability model. The third column reports the change in the probability of stunting predicted by the change in the mean value of the row-specific predictor. The figures in the last colum indicate the change predicted by the shift in the estimated coefficients. The sum of the changes are reported as "Total," and below these we calculate the changes predicted by the three maternal education dummies (Mo.Ed.), the two paternal education dummies (Fa.Ed.), and the variables that measure the availability of health-related amenities in the village (Health). See the text for details.

we estimate that the proportion of children categorized as "wasted" (i.e., whose weight given height is such that the z-score is below -2) decreased by approximately 3 percentage points in rural areas and by 5 percentage points in urban areas. The results for height-for-age, a measure of long-term nutritional status, are instead mixed, and we only find improvements in urban areas, where the proportion of "stunted" children (i.e., with a z-score below -2) decreased by approximately 3 percentage points between 1992–93 and 1998–99. However, we also document that these overall changes hide large differences between genders and across different geographical areas. In fact, we find that nutritional status improved substantially more for boys than for girls. Notably, our results show a small decline in height performances for girls in rural areas, while for boys we observe an almost identical but positive change. We also document the existence of apparent geographical differences in these changes: the gender differences in the changes in nutritional status appear to be driven by rural areas of North India areas, where the existence of widespread son preference has been documented by an immense body of research.

We also make a first attempt at evaluating the determinants of the changes over time, studying the relation between changes in stunting and changes over time in a list of economic and demographic variables that should be strongly associated with child growth performance. Overall, we find that changes over time in the level of the predictors leave much of the actual change in the distribution of height-for-age z-scores unexplained, especially in rural areas. Oaxaca decompositions of the probability of stunting confirm that, for both genders and across all of India, most of the change in anthropometric performances is explained by changes in the relationship between z-scores and the predictors rather than by changes in the predictors themselves. However, a detailed analysis of the patterns of the changes in the coefficients does not point to a simple explanation for the emerging gender differences that we document.

The unequal improvements for boys and girls are all the more difficult to explain because the NFHS suggests that such factors as fertility behavior, women schooling, and female labor force participation changed in ways that would suggest a generalized increase in the relative standing of women in the economy and in the Indian society more generally. At the same time, the relatively large samples available in the surveys we have used, the results 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. Unfortunately, our data set does not allow us to examine directly the possible effect on gender bias in intrahousehold allocation of resources of factors such as male versus female wages, dowries and marriage expenditures, or more generally expected returns to boys versus girls.

It would be useful to corroborate our results 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 was being conducted in the field at the time this article was being written. Another obvious next step would be to study directly the pathways from poverty reduction to child nutrition outcomes, looking in particular at the possible impact of the recent wave of economic liberalization.

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