Child Nutrition and Son Preference in India during the nineties*

Alessandro Tarozzi
Duke University

William H. Dow
University of North Carolina Chapel Hill

March 2004

1 Introduction

Discrimination against women, and especially against young girls, is a well-known and still widespread reality in India, particularly in the Northern States (see, among the countless others, Basu (1992), Dasgupta (1993), Miller (1981), Murthi, Guio and Drèze (1996)). A noted manifestation of this discrimination has been the unnaturally high ratio of men to women in these areas, at all ages. Previous research suggests that proximate pathways through which sex ratios are affected include sex differences in domains such as abortion, infanticide, child health care, and child nutrition.

Several studies from such different disciplines as Anthropology, Economics and Sociology, have found that son preference is particularly strong in areas where the cultural, social, and economic role of women in the society and/or within the household is weaker. So, for example, excess female mortality among young girls is more common in areas where the role of women as bread-earners is smaller, where dowries are more common, or where bequests tend to favour sons over daughters (see Dasgupta (1993, Ch. 11) for an overview).

Many of the factors associated with son preference have economic content.[1] For this reason, economists have explored the possibility that in resource constrained households, even if the welfare of boys and girls enter equally into the parents' utility function, utility maximization will entail a certain degree of preferential treatment for boys. Son preference, in this sense, might therefore be an unfortunate but rational response to unequal economic "returns" to boys and girls. In a seminal paper, Rosenzweig and Schultz (1982), using census and survey data, suggested that parents allocate more resources to boys—therefore improving

*****VERY** preliminary. Comments and suggestions welcome. Please do not quote without permission of the authors.** We thank Jean Drèze for helpful discussions. All errors are ours. Address: Tarozzi: Department of Economics - Duke University, PO Box 90097, Durham, NC 27708, E-mail: taroz@econ.duke.edu. Dow, CB#7411, School of Public Health, University of North Carolina Chapel Hill, North Carolina 27599-7411. E-mail: will_dow@unc.edu.

[1] One exception, for example, might be the importance of males in performing certain religious rituals, which is especially common in North India.

their survival rates relative to girls—when employment opportunities become more biased in favor of men. Behrman (1988), using data from a small number of Indian villages, finds that pro-male bias in intrahousehold allocation of resources is starker during the lean season, when resource constraints are more likely to bind. Garg and Murdoch (1998) suggest that pro-male bias, when associated with resource and credit constraints, may imply better outcomes for children with more females among their siblings. Their empirical analysis support this hypothesis for children from Ghana. Das Gupta (1987) has shown that child mortality differentials are particularly stark for high birth order girls in Punjab. Muhuri and Preston (1991) found similar results in Matlab (Bangladesh). More recently, Pande (2003), using data from the 1992-93 round of the Indian National Family Health Survey, found pro-male bias in immunization rates and prevalence of severe stunting.

The fact that resource constraints—coupled with pro-male bias in economic opportunities—appear to provide an economic "rationale" for the existence of son preference, might lead to expect *reductions* of son preference 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 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 has contributed to a *worsening* of gender bias, as the desired number of sons has decreased less quickly than the desired total number of children (Das Gupta and Bhat, 1995). Overall, these observations lead to ambiguous predictions on the relation between son preference and economic development, which is further complicated by the observations that growth might also be associated with an increase in gender disparities in the labor market.

India experienced several years of fast economic growth in the nineties, following an important and still ongoing process of economic liberalization that started in Summer 1991, after a balance of payment crisis. According to several observers, this process has been associated with a large reduction in poverty (see, for example, Deaton and Drèze (2002)). The first goal of this paper is to establish some stylized facts on the changes in boy versus girl nutritional status during this period. For this purpose, we use data from the two waves of the Indian National Family and Health Survey, carried out in 1992-93 and 1998-99.

2 Data

The main source of our data are the two waves of the Indian National Family Health Survey (NFHS) that are available at the time of writing. The first wave (NFHS-I) was carried out in 1992-93, shortly after the start of the economic reforms in 1991, while the second wave (NFHS-II) was completed in 1998-99. These surveys contain a wealth of information on health, fertility and other family issues for a sample of ever married women aged 13-49 years (15-49 for the later wave). The two samples have been drawn independently from the population of interest, so that the complete dataset is *not* a true panel, but rather two independent cross-sections. In Table 1 we report a list of summary statistics calculated separately for the rural and the urban sector. Many indicators suggest wide changes are taking place. Fertility is declining both in cities and in rural areas. Family size, both actual and desired, is declining, and the use of contraceptives has become more common. In urban areas, the proportion of women *not* practicing birth control declined from 52 to 45 percent. In rural areas the proportion declined from 65 to 58 percent. The proportion of women desiring no more than three children increased from 80 to 86 percent in urban areas, and from 65 to 73 in rural areas. Interestingly, we do *not* find evidence that the number of desired boys is not declining as fast as the number of desired children. Actually, we find that the proportion of desired girls is *higher* in 98-99 than in 92-93, even if the proportion is still below one half. Female illiteracy is declining but still widespread, especially in the countryside, where approximately 70 percent of interviewed women are illiterate. The proportion is 50 percent lower in urban areas, where the fraction of women with at least a secondary school diploma almost doubled, from 10.6 percent in 1992-93 to 18.3 in NFHS-II. Female labor force participation is also on the rise, especially in rural areas where, however, less that two thirds of women work for a salary.

In the central part of Table 1 we also report some first evidence on the existence of gender bias. Because each woman was asked to report a complete birth history, we can calculate a measure of male versus female mortality rates, as in Rosenzweig and Schultz (1982). So, for each interviewed woman who reported giving birth to at least one male and at least one female, we compute the difference between the proportion of surviving males and that of surviving females. We also calculate separate means for North, East, Center, South, and for all Union Territories (see the table for a complete list of the states entering each group). In line with the existing literature, we find that gender bias is concentrated in the Northern states, especially in rural areas, where the fraction of surviving boys is approximately 3 percentage points higher than the fraction of surviving girls, in 1992-93. There is a slight decline in the differential mortality gap between NFHS-I and NFHS-II. In urban areas, the gap is still positive, but extremely small, and not statistically different from zero using standard significance level. The only other region where we find a statistically significant male advantage is rural Center, in 1992-93, but not in the following wave. Note also that the *female* advantage apparent in the South in early nineties is no longer present in the later period examined by the NFHS. This is consistent with the evidence from other recent studies that suggest that the “North-South” divide in gender bias is becoming less clear-cut over time (see, for example, Agnihotri (2003)). The overall change for all India suggest a *slightly increasing* male advantage, both in cities, where the gap remains however very small and not statistically significant, and in rural areas, where boys’ survival rates are approximately one percent higher than those for females.

Both NFHS waves also recorded many information on the respondents’ children, including anthropometric indicators. Namely, height and weight were measured for children below age 4 (in NFHS-I), or below age 3 (in NFHS-II). Because of lack of appropriate measuring tools, NFHS-I did not measure height in Andhra Pradesh, West Bengal, Himachal Pradesh, Madhya Pradesh, and Tamil Nadu, while NFHS-II contains data for all India. In the next section we provide a more thorough description of the measures of nutritional status that we will use throughout the paper.

3 Child Nutritional Status

For each child, the surveys also report measures of nutritional status relative to the CDC growth charts for American children, an international reference population recommended by the World Health Organization (Dibley et al., 1987a; 1987b). These reference growth charts have been widely used as an international standard for cross country anthropometric comparisons. Their use is based on empirical findings supporting the hypothesis that well-nourished children in different population groups follow very similar growth patterns (Martorell and Habicht, 1986). Agarwal et al. (1991) showed that these charts describe adequately the growth process of well-fed Indian children. If weight (given sex and age or height) or height (given age and sex) in the reference population are approximately normally distributed, the relative nutritional status of child i can be evaluated looking at his/her *z-scores*, which are calculated as $(m_{ig} - m_g)/sd_g$, where m_{ig} is weight or height for a specific child in group g (defined by sex and either age or height), and m_g and sd_g are respectively the mean (or median) and the standard deviation of the indicator for children within the same group in the reference population. So, for example, if in the reference population weight-for-height is normally distributed, a boy with a z-score below -1.65 would have a weight below that of 95 percent of boys in the reference population with the same height. Children are frequently deemed to be *stunted* if their height-for-age z-score is below -2, and *wasted* if their weight-for-height is below the same rule-of-thumb threshold. We do *not* use the z-scores calculated in NFHS, because such variables present the perplexing feature of changing considerably among children of the same weight, height, and age. This should not be the case, as they should all share the same reference, and therefore should have the same z-score. So, for example, for one-month old boys whose weight is 3.3 kgs the NFHS contain z-scores ranging from -0.68 to -1.89.

For this reason, we calculate z-scores from the raw data on weight and height. The reference we adopt is the latest revision of the pediatric growth charts for American children (CDC, 2000). These revised charts take into account the fact that weight and height are actually *not* normally distributed. So, to retain the interpretation of the z-scores in terms of percentiles of the reference population, they should be calculated in terms of normal transformation of weight-for-age in the reference population. For this purpose, in the revised 2000 growth charts, the CDC used a Box-Cox transformation of the indicators in the reference population. A Box-Cox transformation of a random variable y can be found estimating the parameter λ, called power, that makes the transformation $(y^\lambda - 1)/\lambda$ as close as possible to a normal distribution.[2] One can show that

[2] See Box and Cox, 1964, or Davidson and McKinnon, 1993, Ch. 14). For details on the use of such transformation in the CDC growth charts, see Kuczmarski et. al. (2000).

this implies that z-scores have to be calculated as

$$z_{ig} = \frac{(X_{ig}/M_g)^{\lambda_g} - 1}{S_g \lambda_g}$$

where M_g , S_g , and λ_g are the median, the standard deviation, and the power of the Box-Cox transformation.[3]

4 Nutrition and poverty

That better nutrition is generally associated with more available resources is a well established fact, even if the literature is not unanimous in quantifying the elasticity of nutrient intakes (or nutritional status) with respect to income.[4]

Here we document the existence of a clear relation between poverty and nutritional status using NFHS data matched with poverty estimates obtained from the Indian National Sample Survey (NSS), the large expenditure survey routinely used to monitor poverty in India. For decades, the Planning Commission of the Government of India has regularly published "official" headcount poverty ratios, separately for rural and urban areas of every Indian state and Union Territory. The poverty counts are computed as the fraction of the population living in households with consumption per head below a poverty line. The poverty lines have been calculated to represent the minimum monthly expenditure per head associated, on average, with a sector-specific minimum calorie intake, recommended by the Indian National Institute of Nutrition. The lines are kept constant in real terms by using two different state-specific price indexes: the Consumer Price Index for Agricultural Labourers (CPIAL) for rural areas, and the Consumer Price Index for Industrial Workers (CPIIW) for the urban sector.[5] Expenditure data are collected by the Indian National Sample Survey Organization (NSSO) approximately every five years, from a large sample of Indian households interviewed over a one-year period. Deaton and Tarozzi (2000), and Deaton (2003) criticize the appropriateness of the indexes used to price inflate the lines, and propose alternatives. Here we use the sector and state specific poverty headcount ratios from Deaton (2003).

We use poverty estimates from the 50^{th} and the 55^{th} wave of the NSS. The former round was carried out between July 1993 and June 1994, and the latter from July 1999 to June 2000. Therefore, we can link poverty rates from the 50^{th} (55^{th}) round to child nutritional status from the 1992-93 (1998-99) wave of the NFHS. Both pairs of surveys have been carried out during time frames that do not perfectly overlap, but they are close enough to allow meaningful comparisons.

[3]The values of these parameters necessary for the calculation of the z-scores can be downloaded from the CDC web site, at http://www.cdc.gov/nchs/.

[4]Recent surveys on this topic can be found in Deaton (1997), and Strauss and Thomas (1998). See also Haddad et al. (2003).

[5] For a detailed overview of the issues related to the choice of poverty lines in India see GOI, Planning Commission (1993), or Deaton and Tarozzi (2000).

For each time period, and for each state and sector for which poverty estimates are available, we calculate the average z-score for weight-for-age, height-for-age and weight-for-height. Then we calculate the correlation between sector and sex-specific mean z-scores and the headcount poverty rates. We report the results in Table 2. All correlations are negative, as expected, and most are above .5. Interestingly, the correlations appear to be much stronger in the earlier period than in the later one. Also, in the earlier nineties the correlation is systematically stronger for males than for females, while in the later period the picture is mixed.

In the next section we turn to examine more closely the changes in child nutritional status during the examined period.

5 Empirical evidence on Nutritional status

In this section we study the relative nutritional status of boys versus girls, and its evolution over time. We proceed estimating nonparametric kernel densities of different measures of nutritional status.[6] Most of the densities refer to z-scores, calculated as described above, but in evaluating the evolution of nutritional status over time we also consider raw weight and height, because we do not want our finding to depend on the choice of a normalization. Clearly, a reference is necessary when evaluating anthropometric performance in a cross-section. To highlight boys-girls differences, or changes over time, in most cases we display both the estimated densities and the difference between the corresponding cumulative distribution functions (CDF's). Such differences are very useful to highlight the discrepancy between densities, which are often difficult to read simply by graphical inspection. In all graphs using z-scores we will include vertical lines corresponding to 0 and -2: these two values correspond to the median and to the 2.5 percentile of the distribution in the reference population.

Because NFHS-I did not measure height in Andhra Pradesh, West Bengal, Himachal Pradesh, Madhya Pradesh, and Tamil Nadu, while NFHS-II contains data for all India, one has to be careful in building comparisons over time. Another reason to be cautious is the fact that NFHS-I recorded anthropometric data for children less than four years old, while NFHS-II considered only children less than three years old. Also, when comparing the distribution of different anthropometric indices in NFHS-I, we have to take into account that weight-for-age can be calculated for all India, while weight-for-height and height-for-age cannot.

However, NFHS-I data suggest that nutritional status in the states where height was not recorded is not very different from that in the rest of India, at least judging from weight performance. Similar arguments suggest that the distribution of nutritional indicators is almost identical for less than 3 and less than 4 years old children. Figure 1 shows that weight-for-age z-scores for boys and girls are almost indistinguishable among the different sub-sample.

[6] We use an Epanechnikov and we adopt Silverman robust bandwidth selector (Silverman (1986)).

Now we turn to examining gender differentials in nutritional status in 1992-93. Because we want to compare discrepancies between genders across different indicators, we consider only the states for which we have valid height data. Also, to ensure comparability with the later figures from NFHS-II, we only include children less than 3 years old. The figures for the full sample are very similar, and are available upon request from the author. Figure 1 also displays densities for weight-for-age for the full sample. The graphs in Figure 2 clearly show that the vast majority of children in the sample have a very poor nutritional status. All densities mostly lie to the left of zero, and for weight-for-age and height-for-age the median is around or below -2. However, the graphs do not show large differentials between boys and girls. In particular, *Height performances appear to be better for girls*, as the corresponding density appear to be shifted to the right with respect to the one for boys. However, from the top-left figure one can see that girls appear to be more likely to be severely underweight, as the left tail of the density is "fatter" for girls than for boys.

The discrepancies are more clearly seen if one looks at the differences between *CDF*'s. Figure 3 shows such differences for the three anthropometric indicators. Each line plots $CDF_{boys}(z) - CDF_{girls}(z)$, where z is a given value for the z-score. Because a *CDF* is the probability that a variable is *below* the specified argument, a *negative* difference implies the presence of a *male advantage*. So, in a situation where girl nutritional status—relative to the reference population—is systematically below that of boys, one would expect the differences to be below zero over the relevant range.[7] Clearly, such situation does not characterize the children in this sample. Male advantage is present for very low weight, but not for weight around or above -2, while female have better relative anthropometric indices for every value of z. Overall, these gaps are not large (note the scale of the vertical axis), and in almost no case the difference is above 3 percent in absolute value. The curve for weight-for-height is almost flat at zero, showing a slight female advantage for very low values, and a small male advantage for larger values, where, however, most of the children are (see Figure 2). Weight-for-height is frequently considered an excellent indicator of nutritional status, as it has the advantage of non depending upon age, which is frequently misreported in poor countries, and because it can distinguish between tall but thin children and small but well nourished ones.

Figure 4 reports the differences between *CDF*'s for NFHS-II. To ensure comparability with the graphs in the previous figures, we exclude observations from Andhra Pradesh, West Bengal, Himachal Pradesh, Madhya Pradesh, and Tamil Nadu. The results show a striking change *in favor of males*. The male advantage for weight-for-age is now more pronounced, and it stretches over a wider range. Girls still have an advantage along the upper tail, but the gap in their favor is much less pronounced than in 1992-93. A similar reduction of girl advantage can be seen for height-for-age. As a consequence, the curves for weight-for-height, in 98-99, show a uniform male advantage, so that the distribution of z-scores for males stochastically dominates the distribution for girls.

[7] Technically, this would mean that the distribution for boys first-order stochastically dominate the distribution for girls. Note also that even in this case the differences will be zero for very low and very high values of the argument, as all CDF's are zero if evaluated below the lower bound of the range of the variable, and they are equal to one above the upper bound.

Clearly, these results do not show that girl nutritional status has worsened in absolute terms, but only that it has worsened relative to that of boys. So, we turn now to analyze explicitly changes in nutritional status for boys and girls separately. Each graph in Figure 5 plots the two sex-specific changes over time of the corresponding densities. Because each line is calculated as $CDF_{98-99}(z) - CDF_{92-93}(z)$, improvements in relative nutritional status are represented by *negative* values. Looking at weight-for-age and weight-for-height, there are large improvements for both boys and girls, but the progress is clearly larger for males. For example, the fraction of boys with a weight z-score below -2 dropped by almost 7 percent between 92-93 and 98-99, while the corresponding fraction for girls dropped by only 2 percent. Similarly, there are about 6 percent less boys whose weight-for-age z-score is below -1, but the proportion dropped by 4 percent only for girls. The gender differences are ever starker if one looks at height-for-age, the only measure of nutritional status for which girls showed an advantage over boys in the cross-sectional figures shown before. The top right panel in Figure 5 shows that weight-for-age for females was actually slightly *better* in 92-93 than in 98-99. For boys, instead, there is a small improvement in this dimension too. So, the overall improvement in girl weight-for-height is partly due to higher weight-for-age, and partly to *lower* height-for-age.

Because we do not want these results to be driven by the reference population we choose for the calculation of z-scores, we also evaluate the changes over time of anthropometric indicators using directly height and weight. We use nonparametric regressions to estimate growth charts for weight or height given age, and for weight given height. The use of nonparametric tools allow us to study the regressions without the need of making functional form assumptions. Locally weighted regressions (Fan, 1992) estimate the conditional expectation at a given point x through a simple "local" weighted OLS regression. The regression is "local" in the sense that only observations in a neighborhood of x are included in the regression, and it is weighted since a kernel centered around x is used to attach larger weights to points closer to x.[8]

Another reason to look at the raw data on weight and height is the fact that, even if the choice of reference population is appropriate, z-scores can change systematically with age. It is well known that in areas where poor nutrition is widespread, z-scores frequently decrease during the first months of life, as the negative effect of poor nutrition and health accumulate, and often stabilize only around 18 months of age. In Figure 6 we show that this pattern is also prevalent in India.[9] If the empirical distribution of age is different in 1992-93 and in 1998-99, then, changes in the *distribution* of z-scores might reflect changes in the age distribution rather than changes in nutritional status. Figure 7 shows that the age distribution is, in fact, not identical in the two different waves. A formal chi-square test for the null of equality of the distributions across surveys—adjusted for the presence of clustering—rejects the null even using a 1 percent level. Note

[8] Fan (1992) shows that locally weighted regressions possess better properties than the traditional kernel regression estimator, since they reduce the bias arising in the estimates when the density of the regressor is not flat, a bias which can be particularly large near the boundaries of the regression. Deaton (1997, Ch. 3) provides an intuitive introduction to both estimators. For a more technical treatment see, for example, Pagan and Ullah (1999, Ch. 3).

[9] Here we only show graphs from 1992-93, pooling boys and girls together. Clearly, one could also examine changes in average nutritional status between the two NFHS waves comparing sex differences or changes over time in these graphs.

also the age misreporting is apparent, as the distributions present "fluctuations" that would be difficult to explain otherwise. Interestingly, there is no evidence that age misreporting leads to "peaks" at focal ages like 12, or 24 months of age.

To use the non-normalized data on weight and height, we proceed as follows. First, for both sexes, and each NFHS wave, we estimate nonparametric regressions of weight on age, height on age, and weight on height. The bottom right panel in Figure 8, for example, plots the estimated nonparametric regressions for boys in 1992-93 and 1998-99. Then, we calculate changes over time of each regression, plotted against the corresponding regressor. The three sex-specific pairs of estimated changes are displayed in Figure 8. So, for example, the curve labelled "boys" in the top-left panel represents the vertical distance $\hat{E}\,[\text{weight} \mid \text{age}, 98\text{-}99] - \hat{E}\,[\text{weight} \mid \text{age}, 92\text{-}93]$ where the two estimated conditional expectations are shown in bottom-right figure. By construction, improvements in nutritional status translate into *positive* differences. The information contained in the two left-most figures is consistent with the changes in CDF's displayed before. Weight performances, given age or height, improve for both boys and girls, but the changes are clearly larger for boys for almost all the relevant range. For height-for-age the picture is mixed, and improvements are only apparent for younger children, while the differences are negative for older ones. Note also that here there is no clear male advantage.

6 Regression analysis

In this section we use regression analysis to examine determinants of child nutritional status, emphasizing changes that intervened between the two waves of the NFHS waves. We are particularly interested in studying how child nutrition is affected by economic factors, as we are evaluating a period of time of rapid economic growth. The NFHS does not record household income or consumption, so that we cannot use them to evaluate the association between child nutrition and household resources. For this purpose, we use instead available information on asset ownership. Ownership is recorded for different lists of items in two waves, and to ensure comparability we use only variables that are common to both surveys. In NFHS-II, most of the assets have been used to assign to each household a "low", "medium", or "high" "standard of living index", constructed assigning predetermined scores to different asset ownership indicators (see the NFHS final report for details). We construct an analogous standard of living index including only the variables that are common to both surveys. We use indicators for several housing characteristics, land ownership, and ownership of various durables. The complete description of the variables included in the index, and the scores assigned to each, can be found in the appendix.

As an alternative to this standard of living index, we also use principal component analysis to construct an asset index. Such statistical procedure aims at describing the variance of a set of k variables in terms of k orthogonal factors (or principal components) whose meaning should be suggested by the context. Frequently, the first principal component explains the largest share of the total variance, and it is therefore interpreted as a "factor" that all variables have in common. In the present context, where the original variables measure asset ownership, it is natural to think that the first principal component can be used as a proxy for the household's wealth. We report the details of the calculations in the appendix.

We consider determinants of the probability of *stunting* and *wasting*, and we run separate probit regressions for each sector and wave. The former is defined by a height-for-age z-score below -2, while the latter is present when the z-score for weight-for-height is below the same threshold. All regressions are carried out using the sampling weights provided in the surveys. Most surveys, including the NFHS, first sample *clusters* (typically villages, or urban blocks), and then sample households from the selected clusters. This generally introduces intra-cluster correlation, which may lead to large increases in the standard errors of the estimated statistics. For this reason, all standard errors are corrected to take into account the presence of intra-cluster correlation.[10]

As in the previous section, we only include observations from states for which we have data on both weight and height for both NFHS waves. Table 3 contains summary statistics for child nutritional status and for the regressors, which include child age (and its squared), religion, household size and wealth (proxied by one of the two asset index indicators we have described previously), indicators for mother's work status and parental education, district-wise male and female employment rates, and sibling composition. We also include state dummies, and indicators for twins, and for mothers more than 40, or less than 20 years old. Consistently with the evidence presented in the previous sections, we can observe improvements in nutritional status, with the exception of the proportion of stunted children in the rural sector, which remains basically unchanged. Note also that both proxies for household wealth show marked increases over time, which is again consistent with the reduction in poverty indicated by the NSS estimates over the period.

We report probit estimates, for each sector and wave separately, in Table 4 (for stunting) and Table 5 (for wasting). In these regressions we use the first principal component as proxy for household wealth, but the results do not change substantially if we use the "standard of living index" instead. Because we are interested in evaluating son preference, we have included several interactions between regressors and a dummy equal to one if the child is a girl. Exploratory analysis suggest that including further interactions does not add important insights, and we have chosen a relatively parsimonious model to avoid clutter. The figures displayed in the table represent marginal effects, and can be interpreted as the predicted change in the binary dependent variable associated with a unit change in the regressor, keeping all other regressors equal to the corresponding sample mean. All the coefficients should be interpreted with caution. Because child nutrition, within each household, is likely to be determined jointly with fertility and labor decisions, some of the regressors are likely to be endogenous, so that it is not clear that one can give them a *causal* interpretation. Still, the results should be interesting, as they identify variables that are associated with the

[10] For an overview of the issues involved in the calculation of standard errors when data come from multi-stage surveys see Deaton (1997, Ch. 2).

probability of stunting or wasting.

We first look at the results for the probability of stunting, which are reported in Table 4. Somewhat surprisingly, most regressors do not enter significantly the regression. As expected, older children are significantly more likely to be stunted, as the negative effect of poor nutrition and health accumulate over time. The negative coefficients for the quadratic terms arise because z-scores generally stabilize after 18-24 months of age. None of the indicators for the mother's work status enters the regression with significant coefficients, both for males and for females. Similarly, we do not find a significant association between the probability of stunting and district-wise female and male employment rates. Not surprisingly, parental education does instead enter significantly the regression. Having an illiterate mother is estimated to increase the probability of stunting by 6 to 10 percentage points, while the negative impact of an illiterate father is not as strong, and it is significant only for the rural sector in the later wave. We cannot reject the hypothesis that the impact of parental education for girls is the same as for boys. There is some evidence that parental education is more important in the later survey than in the first one. Interestingly, we cannot reject the hypothesis that sibling composition affects boys and girls in the same way, suggesting that some of the opposite results found in the literature might not be robust to changes in the regression specification.[11] Overall, we do not find strong evidence that the correlates of stunting are different for boys and girls. The hypothesis that *all* female interactions are zero is strongly rejected in both sectors and in both periods, but this might be due to the high power of the test, as the sample sizes are large, and the number of restrictions is large as well. When we test for the statistical significance of subsets of female interactions (see bottom of the panel) we never reject the null of zero coefficients.

The overall picture is not very different when we move to examine the determinants of wasting, in Table 5. Here, if anything, the estimated coefficients appear to be even more erratic than in the previous regressions. This might be due to the fact that weight-for-height is much more sensitive than height-for-age to short term changes in nutritional status, and our list of included regressors might be inadequate to take such short term elements into account. This is consistent with the large decrease in the pseudo-R^2, which in the probit regressions for stunting was in the range of 0.12-0.15, and is now in the range of 0.3-0.5.

[11] For example, Pande (2003), using a logit model and a different set of regressors, finds that girls are harmed more than boys by the presence in the household of more older sisters.

7 References

Agarval, K. N. et al., 1991. Growth performance of affluent Indian children (under-fives): growth standards for Indian children. Nutrition Foundation of India, New Delhi.

Agninhotri, S B, 2003, "Survival of the girl child", *Economic and Political Weekly*, October 11, 4351-4360.

Basu, Alaka Malvade, 1992. Culture, the status of women and demographic behaviour. Oxford Clarendon Press.

Behrman, J. R., 1988, "Intrahousehold allocation of nutrients in rural India: are boys favored? Do parents exhibit inequality aversion?", Oxford Economic Papers, 40, 32-54.

Box, G. E. P. and D. R. Cox, 1964. An analysis of transformations. Journal of the Royal Statistical Society, Series B, 26, 211-243.

Das Gupta, Monica, 1987. "Selective discrimination against female children in rural Punjab", Population and Development Review, 13 (1).

Das Gupta, Monica, and Mari Bhat, 1995. "Intensified gender bias in India: a consequence of fertility decline". WP no. 95.02, Harvard Center for Population and Development Studies.

Dasgupta, Partha, 1993. An inquiry into well-being and destitution, Oxford University Press.

Davidson, Russell, and James G. MacKinnon, 1993. Estimation and inference in econometrics. Oxford University Press, Oxford, UK.

Deaton, Angus, 1997. The analysis of household surveys. Johns Hopkins University Press, Baltimore, MD.

Deaton, Angus S. (2003a), "Prices and poverty in India, 1987-2000", *Economic and Political Weekly*, January 25, 362-368.

Deaton, Angus S., and Jean Drèze (2002), "Poverty and inequality in India, a re-examination", *Economic and Political Weekly*, September 7, 3729-3748.

Deaton, Angus S., and Alessandro Tarozzi (2000), "Prices and poverty in India," (Research Program in Development Studies Working Paper no. 196, Princeton University).

Dibley, M.J., J.B. Goldsby, N.W. Staehling, and F.L. Trowbridge. 1987a. Development of normalized curves for the international growth reference: Historical and technical considerations. American Journal of Clinical Nutrition 46(5): 736–748.

Dibley, M.J., N.W. Staehling, P. Neiburg, and F.L. Trowbridge. 1987b. Interpretation of z-score anthropometric indicators derived from the international growth reference. American Journal of Clinical Nutrition 46(5): 749–762.

Fan, Jianquing, 1992. Design-adaptive nonparametric regression. Journal of the American Economic Association 87, 998-1004.

Filmer, Deon, and Lant Pritchett, 2001. Estimating wealth effects without expenditure data - or tears: an application to educational enrollments in states of India. Demography 38 (1), 115-132.

Garg A and J Morduch, 1998, "Sibling rivalry and the gender gap: evidence from child health outcomes in Ghana", Journal of Population Economics, 11, 471-493.

Haddad, Alderman et al (2003), World Bank Economic Review.

Kuczmarski et. al., 2000. CDC Growth charts: United States. Advance Data 314, Center for Disease Control and Prevention, National Center for Health Statistics.

Lindeman, Richard H., Peter F. Merenda and Ruth Z. Gold, 1980. Introduction to bivariate and multivariate analysis. Scott, Foresman and Co.

Martorell, R. and J. P. Habicht, 1986. Growth in early childhood in developing countries. In Falkner F., and Tanner J. M. (eds.), Human growth: a comprehensive treatise, Vol. 3, Plenum Press, New York.

Miller, Barbara, 1981. The endangered sex: neglect of female children in rural North India, Cornell University Press.

Murthi, Mamta, Anne-Catherine Guio, and Jean Drèze, 1996. "Mortality, fertility and gender bias in India: a district-level analysis". In: Drèze J. and Sen A. (eds.), Indian development: selected regional perspectives, Oxford University Press.

Muhuri, Pradip K, and Samuel H Preston, 1991. "Effects of family composition on mortality differentials by sex among children in Matlab, Bangladesh". Population and Development Review, 17 3).

Pagan, Adrian and Aman Ullah, 1999. Nonparametric econometrics. Cambridge University Press, Cambridge, UK.

Pande, Rohini, 2003. "Selective gender differences in childhood nutrition and immunization in rural India: the role of siblings". Demography, 40 (3).

Rosenzweig, M. R. and Paul T. Schultz, 1982, "Market opportunities, genetic endowments, and intrafamily resource distribution: child survival in rural India". American Economic Review, 72 (4), 803-815.

Strauss and Thomas (1998), Journal of Economic Literature.

8 Appendix

8.1 Construction of the asset index

The two waves of the NFHS do not contain data on income or expenditure, but they contain information on ownership for a large number of assets. The list of assets is not identical between the two waves, and to ensure comparability we only use those that are common to both surveys. In the 1998-99 wave of the NFHS, most of the assets have been used to assign to each household a "low", "medium", or "high" "standard of living index", constructed assigning predetermined scores to different asset ownership indicators (see the NFHS final report for details). We construct an analogous asset index indicator, but we use only variables that are common to both surveys.[12] The following table lists the assets included in the index, and the assigned scores. The scores are identical to the ones used for the same items to construct the standard of living index.

Table A1 - Scores for the calculation of standard of living index.

House type:	4 for pucca, 2 for semi-pucca, 0 for kachha.
Toilet facility:	4 for own flush toilet, 2 for public or shared flush toilet or own pit toilet, 1 for shared or public pit toilet, 0 for no facility.
Source of lighting:	2 for electricity, 1 for kerosene, gas, or oil, 0 for other source of lighting.
Main fuel for cooking:	2 for electricity, liquid petroleum gas, or biogas, 1 for coal, charcoal, or kerosene, 0 for other fuel.
Source of drinking water:	2 for pipe, hand pump, or well in residence/yard/plot, 1 for public tap, hand pump, or well, 0 for other water source
Separate room for cooking:	1 for yes, 0 for no.
Ownership of house:	2 for yes, 0 for no.
Ownership of agricultural land:	4 for 5 acres or more, 3 for 2.0–4.9 acres, 2 for less than 2 acres or acreage not known, 0 for no agricultural land.
Ownership of irrigated land:	2 if household owns at least some irrigated land, 0 for no irrigated land.
Ownership of livestock:	2 if owns livestock, 0 if does not own livestock.
Ownership of durable goods:	4 for a car, 3 each for a moped/scooter/motorcycle, refrigerator, or television, 2 each for a bicycle, electric fan, radio/transistor, sewing machine, 1 for clock/watch.

There are 181,048 households overall, 1768 of which have missing information for at least one asset, and in only 70 households more than 1 value is missing. In case of missing information on an asset, we impute a value equal to the median within the same district. We use the median rather than the mean because most indicators are binary variables, and using means would produce non-binary values that would be difficult to interpret within a principal component framework. We set the asset index to missing for the handful of observations for which more than 8 asset variables are missing.

An alternative—based on the same principle—is to construct a proxy for household wealth by using principal component analysis. Such statistical procedure aims at describing the variance of a set of k variables in terms of k orthogonal factors (or principal components) whose meaning should be suggested by the context.[13] Frequently, the first principal component explains the largest share of the total variance, and it is therefore interpreted as a "factor" that all variables have in common. In the present context, where the original variables measure asset ownership, it is natural to think that the first principal component can be used as a proxy for the household's wealth.[14]

[12] We also exclude information on assets for a handful of items for which information were not collected in the first states surveyed in 1992-93.

[13] For details see Lindeman, Merenda and Gold (1980, Ch. 5 and 8).

[14] Filmer and Pritchett (2001), using NFHS data, show that the asset index gives satisfactory results when used as a proxy for long-run economic status, and they also argue that such index might be a better measure than current consumption or income to evaluate household wealth.

Let X_{ih} be the value of the i-th asset indicator for household h, and let X_i and σ_i indicate the corresponding mean and standard deviation in the sample. Then the first principal component (the asset index) for household h can be written as

$$a_h = \sum_{i=1}^{k} w_i z_{ih}$$

where w_i is the coefficient assigned to the i-th asset, and $z_{ih} = \sigma_i^{-1}(X_{ih} - X_i)$. Therefore, a unit increase in the value of asset i implies a change in the asset index equal to $w_i \sigma_i^{-1}$. When X_{ih} is a dummy variable, the "marginal effects" indicates how the index changes when the household owns the relevant asset.

We construct the asset index pooling together data on asset ownership from both NFHS waves, and using data from all households surveyed, including those without women eligible for an interview. We also calculate separate asset indexes for the rural and the urban sector, as sometimes certain assets seem to indicate a high standard of living in one sector, and a low standard of living in the other. For example, owning livestock, or agricultural land, affects negatively the asset index in the urban sector, while it affects the index positively in rural areas. This is probably explained by the fact that families mostly involved in agriculture are likely to be relatively poorer in cities.

In case of missing asset variables, we proceed as for the standard of living index, replacing missing observations with district medians, and leaving the first principal component as missing if there are more than 8 missing asset indicators. No replacement is necessary for more than 99 percent of the sample. The following table shows the marginal effects for all the included asset indicators, calculated for the rural and urban sector separately, and for all India pooled together.

Table A2 - Marginal effects of asset ownership indicators on the asset index

	All India	Urban	Rural
owns television	0.636	0.622	0.817
owns radio	0.413	0.416	0.480
n. rooms	0.082	0.102	0.120
separate room as kitchen	0.342	0.453	0.380
land cultivated (acres)	0.000	0.000	0.001
land irrigated (acres)	0.000	0.000	0.001
owns livestock	-0.274	-0.193	-0.013
owns clock watch	0.501	0.714	0.537
owns fan	0.590	0.666	0.700
owns sewing machine	0.509	0.455	0.621
main cooking fuel is electr./LPG/Kerosene	0.636	0.544	0.905
drinking water piped into residence	0.572	0.480	0.601
drinking surface/rain water	-0.259	-0.524	-0.141
owns flush toilet	0.659	0.542	0.876
has electricity	0.519	0.789	0.536
no toilet facility	-0.539	-0.615	-0.550
house is pucca (all high quality materials)	0.583	0.535	0.691
house if katcha (all low quality materials)	-0.468	-0.676	-0.458
owns refrigerator	0.760	0.579	1.192
owns bicycle	0.199	0.220	0.263
owns motorcycle	0.677	0.532	0.969
owns car	0.812	0.570	1.312

Source: Author's calculations from NFHS-I,II, 1992-93, 1998-99, All India.

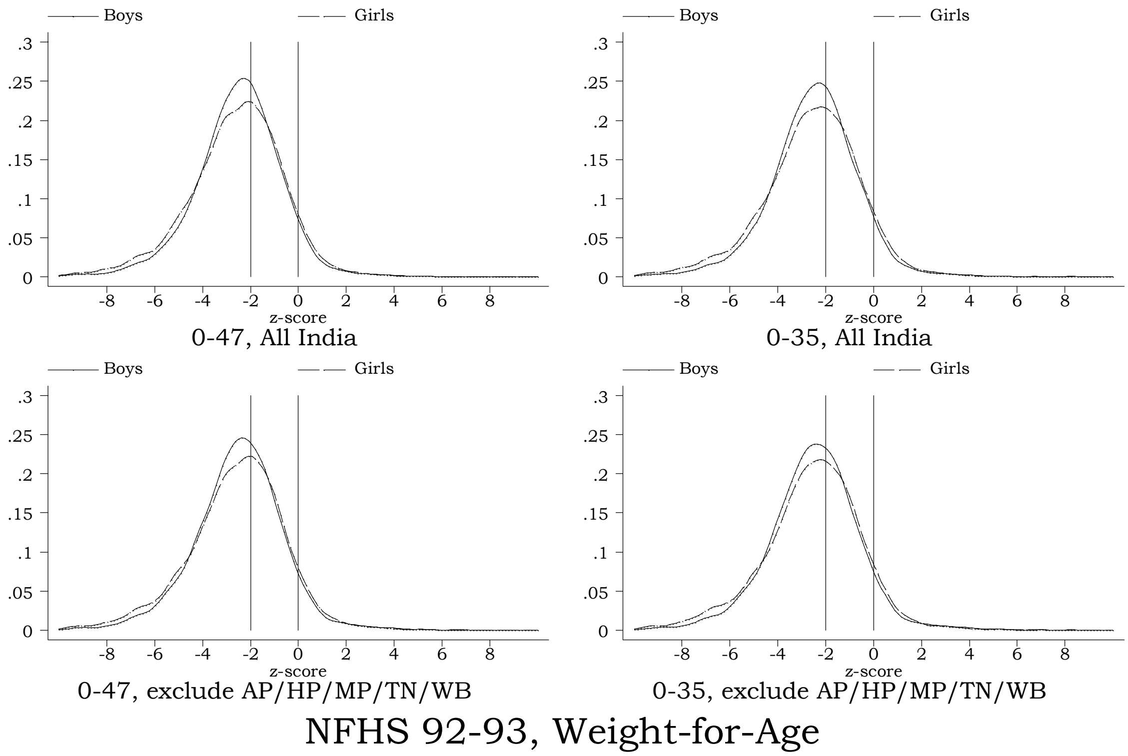

Figure 1: Nonparametric Kernel densities of z-scores for weight-for-age. Source: author's calculations.

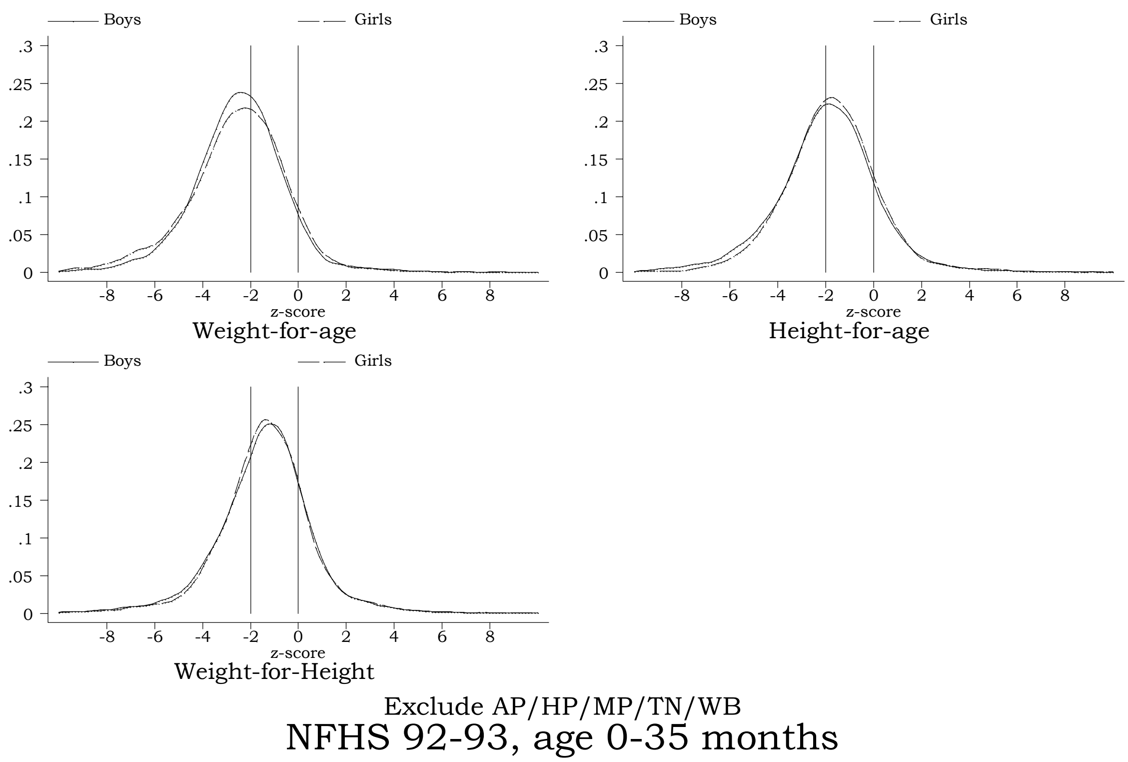

Figure 2: Nonparametric Kernel densities of z-scores. Source: author's calculations.

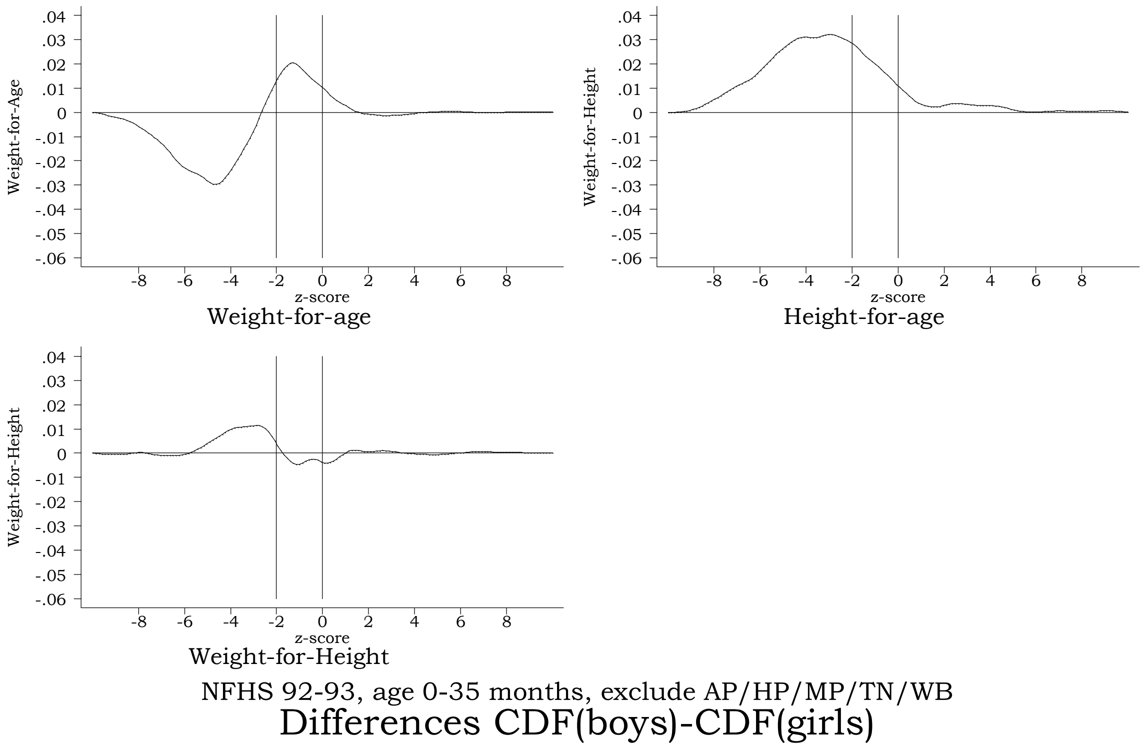

Figure 3: Differences of cumulative distribution functions, estimated by integration of nonparametric kernel densities of z-scores. Source: author's calculations.

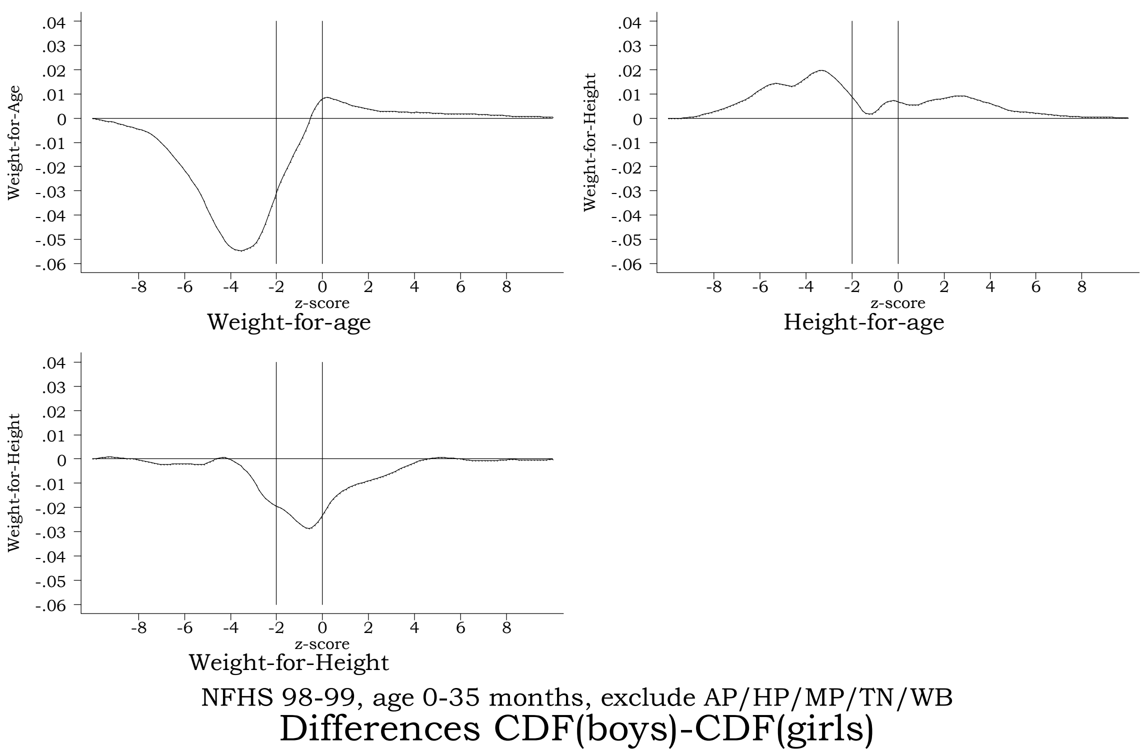

Figure 4: Differences of cumulative distribution functions, estimated by integration of nonparametric kernel densities of z-scores. Source: author's calculations.

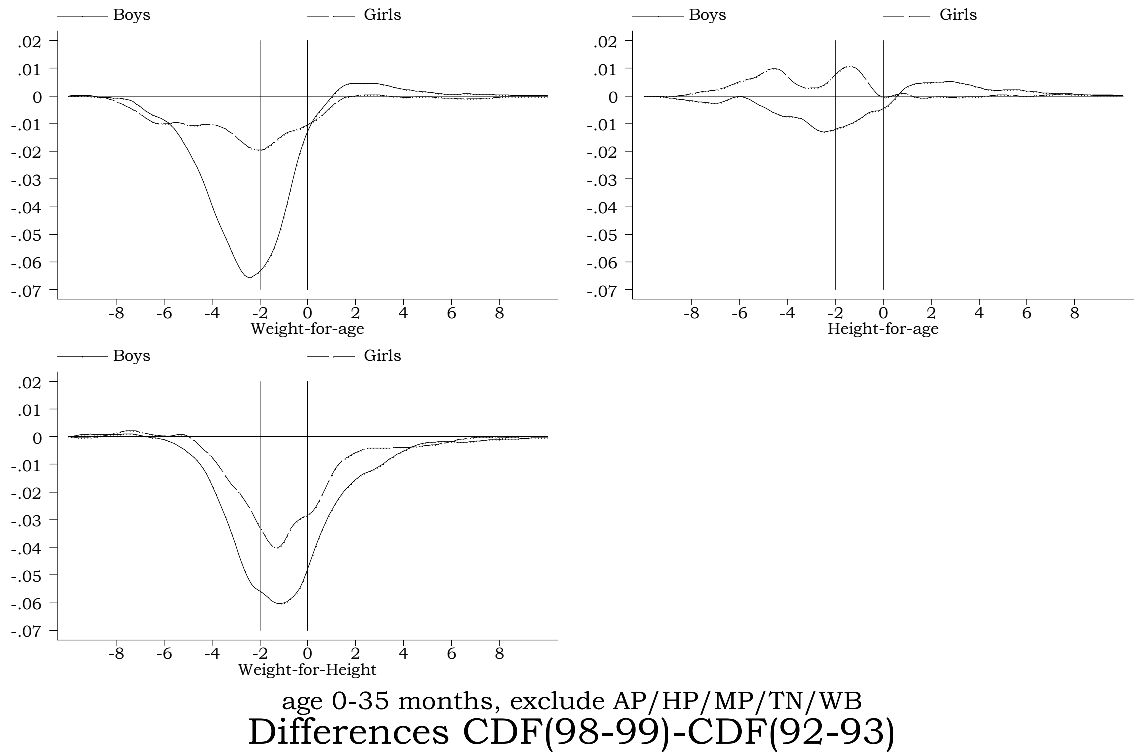

Figure 5: Differences of cumulative distribution functions, estimated by integration of nonparametric kernel densities of z-scores. Source: author's calculations.

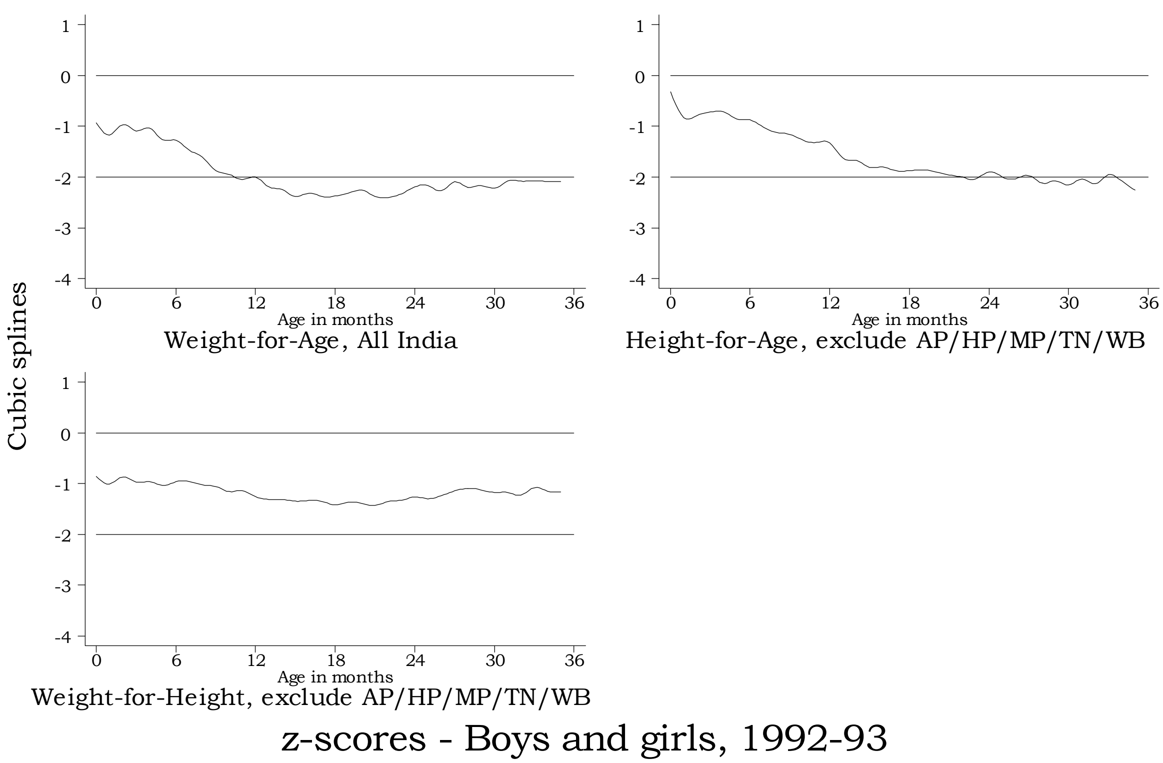

Figure 6: Source: author's calcuations.

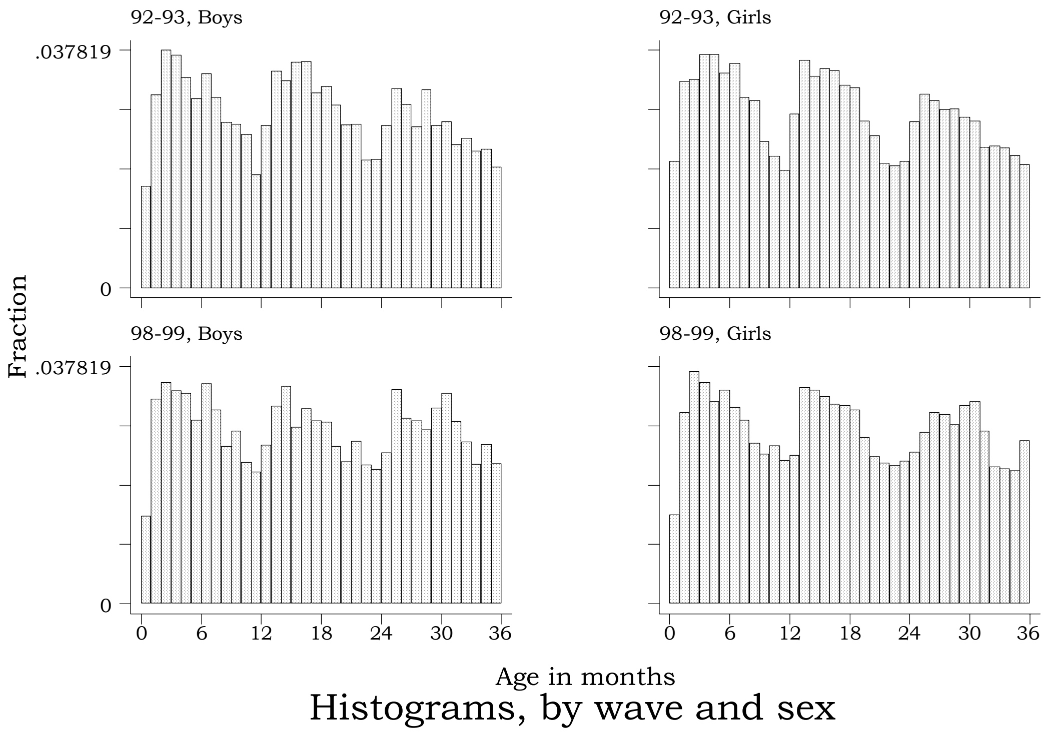

Figure 7: Empirical distribution of reported age for children less than 3 years old. By NFHS wave and sex. Source: Author's calculations.

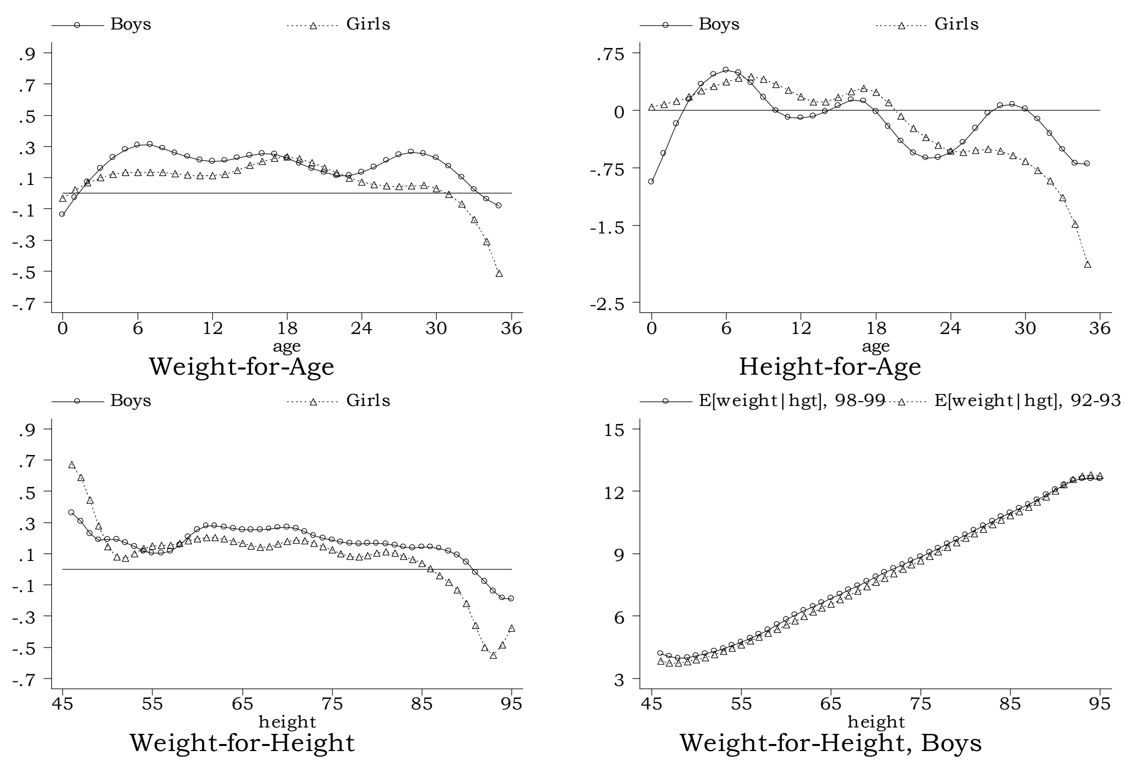

Figure 8: Changes over time of nonparametric regressions of nutritional status.

Table 1 - Summary statistics - Women

	1992-93 (NFHS-I)		1998-99 (NFHS-II)	
No. of households	88562		92486	
No. of ever married women age 15-49	89777		90303	
No. of ever married age 13-15	271		0	
% living in rural areas (weighted)	73.8		73.8	
Means (Weighted)	**Urban**	**Rural**	**Urban**	**Rural**
Family - Fertility				
Age at first marriage	17.9	16.2	18.2	16.4
Household size	6.73	7.24	6.48	6.93
# children below age 5	0.91	1.14	0.81	1.03
Not usign any contraceptive	51.93	65.03	45.46	58.07
Contraceptive: Female sterilization	28.62	24.9	33.72	31.41
Contraceptive: Pill	1.82	0.89	2.53	1.78
Contraceptive: Condom	5.47	1.17	6.75	1.5
% desiring 3 children or less	80.32	65.07	85.56	73.05
% desiring 1 male or less*	57.83	35.81	66.75	45.2
% desiring 1 female or less*	85.43	77.44	89.14	80.54
Desired % of females*	43.4	40.12	44.25	41.52
Male/Female differential mortality**	-0.0007 (.0031)	**0.0107 (.0023)**	0.0017 (.0027)	**0.0115 (.0021)**
North	.0073 (.005)	**.0292 (.0041)**	.0038 (.0047)	**.0243 (.0036)**
Center	-.002 (.0057)	**.0121 (.0058)**	.005 (.0049)	.0089 (.0047)
East	-.0017 (.0091)	.006 (.0047)	.0067 (.0069)	.0062 (.0038)
South	-.0076 (.0062)	**-.01 (.0043)**	-.0069 (.0056)	.0043 (.0049)
Union Territories	-.001 (.0084)	**-.0171 (.0059)**	-.0093 (.0091)	-.0095 (.0059)
Education and Labor Force Participation				
Working	21.25	37.37	24.18	42.05
of which: working for salary	89.23	60.32	88.87	62.66
% illiterate	36.75	72.4	33.16	66.9
% higher secondary completed or higher	10.59	0.81	18.3	2.76
Partner: % illiterate	17.12	40.65	13.45	34.07
Partner: % higher sec. completed or higher	20.8	4.83	33.61	12.07

Source: Author's calculations from NFHS-I and II. All means and proportions calculated using sampling weights.
*Calculated using only responses for which (desired boys+desired girls)=tot. desired children
**Calculated as (sons alive/sons ever born)-(daughters alive/daughters ever born). By construction, this is calculated only for women who had at least one birth of both sexes. The standard errors take into account clustering.
North: Delhi, Uttar Pradesh, Rajasthan, Punjab, Jammu, Himachal Pradesh; Center: Gujarat, Madhya Pradesh, Maharastra; East: Assam, Bihar, Orissa, West Bengal; South: Andhra Pradesh, Kerala, Karnataka, Tamil Nadu.

Table 2 - Correlations between poverty rates and mean z-scores

1992-94	Rural Sector		Urban Sector	
	Males	**Females**	**Males**	**Females**
Weight-for-age	-0.6489	-0.6253	-0.7865	-0.6452
Height-for-age	-0.6424	-0.6857	-0.6368	-0.5486
Weight-for-height	-0.5517	-0.3648	-0.6442	-0.4269

1998-00	Rural Sector		Urban Sector	
	Males	**Females**	**Males**	**Females**
Weight-for-age	-0.6063	-0.514	-0.4629	-0.4883
Height-for-age	-0.2511	-0.1167	-0.2817	-0.3468
Weight-for-height	-0.3704	-0.4527	-0.2968	-0.241

Source: Author' calculations for z-scores, Deaton (2003) for poverty rates.
Each correlation is calculated using one observation per state.

Table 3 - Summary statistics - Variables included in regressions

	1992-93 (NFHS-I)					1998-99 (NFHS-II)				
	Urban			Rural		Urban			Rural	
obs.	6324		#	22146		5242		#	20205	
Means (Weighted)										
Height-for-age z-score below -2	0.37	(0.483)		0.456	(0.498)	0.344	(0.475)		0.465	(0.499)
Weight-for-age z-score below -3	0.483	(0.5)		0.579	(0.494)	0.439	(0.496)		0.542	(0.498)
Weight-for-Height z-score below -2	0.274	(0.446)		0.32	(0.466)	0.239	(0.426)		0.279	(0.448)
Age	17.187	(10.112)		16.718	(10.067)	17.328	(10.168)		17.028	(10.319)
Age squared	397.617	(366.596)		380.841	(359.667)	403.614	(368.189)		396.421	(369.78)
Female	0.482	(0.5)		0.493	(0.5)	0.479	(0.5)		0.475	(0.499)
Asset Index (first principal component)*	-0.576	(2.591)		-0.136	(2.123)	0.052	(2.308)		0.241	(2.342)
Standard of Living Index*	18.76	(8.934)		11.976	(6.524)	21.252	(8.524)		13.887	(7.62)
Muslim	0.172	(0.377)		0.119	(0.323)	0.208	(0.406)		0.138	(0.345)
Christian	0.075	(0.264)		0.074	(0.262)	0.09	(0.286)		0.08	(0.271)
Sikh	0.03	(0.169)		0.041	(0.197)	0.027	(0.162)		0.028	(0.164)
Other Religion	0.023	(0.149)		0.018	(0.135)	0.026	(0.158)		0.024	(0.153)
Mother does not work	0.866	(0.341)		0.709	(0.454)	0.839	(0.368)		0.665	(0.472)
Mother works away from home	0.088	(0.283)		0.253	(0.435)	0.1	(0.3)		0.287	(0.452)
Female employment rate in household district	20.201	(12.499)		21.693	(17.144)	24.203	(14.472)		25.625	(17.642)
Male employment rate in household district	82.797	(4.756)		84.026	(5.459)	82.267	(4.541)		84.009	(4.856)
Mother less than 20 years old	0.065	(0.246)		0.104	(0.306)	0.067	(0.25)		0.106	(0.308)
Mother more than 40 years old	0.011	(0.102)		0.016	(0.126)	0.007	(0.085)		0.013	(0.112)
Mother has no formal education	0.379	(0.485)		0.663	(0.473)	0.317	(0.465)		0.619	(0.486)
Father has no formal education	0.181	(0.385)		0.352	(0.478)	0.13	(0.336)		0.31	(0.462)
Mother has at least secondary school diploma	0.11	(0.313)		0.012	(0.108)	0.212	(0.409)		0.049	(0.215)
Father has at least secondary school diploma	0.175	(0.38)		0.055	(0.228)	0.297	(0.457)		0.144	(0.351)
Household size	7.47	(3.732)		8.17	(4.052)	7.375	(3.65)		7.836	(3.852)
Child is twin	0.014	(0.118)		0.009	(0.094)	0.008	(0.089)		0.011	(0.104)
# of older brothers	0.671	(0.94)		0.81	(1.03)	0.584	(0.846)		0.775	(1.003)
# of older sisters	0.753	(1.049)		0.884	(1.127)	0.702	(1.021)		0.905	(1.12)

Source: Author's calculations from NFHS-I and II. All means and proportions calculated using sampling weights. Standard errors in parenthesis. Does not include Andhra Pradesh, Himachal Pradesh, Madhya Pradesh, Tamil Nadu, and West Bengal.

* See the appendix for a description.

Table 4 - Marginal Effects on the probability of stunting - Probit estimates

	1992-93 (NFHS-I)				1998-99 (NFHS-II)			
	Urban		Rural		Urban		Rural	
	coeff.	t-ratio	coeff.	t-ratio	coeff.	t-ratio	coeff.	t-ratio
Age	0.025	6.27	0.033	15.31	0.032	8.86	0.038	17.75
Age squared	0.000	-2.05	0.000	-7.14	0.000	-4.91	-0.001	-8.88
Asset index (first pr. Comp.)*	-0.027	-3.73	-0.030	-5.91	-0.052	-7.38	-0.028	-6.51
Asset index*Female	0.001	0.1	0.017	2.9	0.007	0.76	0.009	1.53
Muslim	-0.004	-0.16	0.035	1.8	0.023	1	-0.017	-0.9
Christian	-0.033	-0.61	-0.092	-1.96	-0.040	-0.7	-0.057	-1.1
Sikh	-0.071	-1.2	-0.128	-3.54	-0.003	-0.05	-0.113	-2.39
Other Religion (Hindu omitted)	0.094	1.66	0.037	0.68	-0.019	-0.35	-0.090	-1.46
Mother less than 20 years old	0.008	0.21	0.040	2.29	-0.024	-0.62	0.048	2.64
Mother more than 40 years old	-0.058	-0.6	-0.058	-1.39	0.013	0.11	-0.011	-0.24
Mother does not work	0.032	0.56	-0.029	-0.79	-0.085	-1.54	-0.263	-1.3
Mother does not work*Female	-0.061	-0.73	-0.008	-0.16	0.026	0.33	0.008	0.23
Mother works away from home	0.081	1.16	-0.109	-0.63	-0.087	-1.5	0.063	1.25
Mother works away from home*Female	-0.154	-1.7	-0.025	-0.66	0.070	0.73	0.031	0.88
Child is a female	-0.128	-0.35	0.033	0.66	0.002	1.47	0.057	1.09
Female employment rate	0.002	1.71	0.000	-0.19	-0.001	-0.9	0.000	-0.66
Female employment rate*Female	0.000	0.01	0.000	0.76	-0.003	-0.85	-0.001	-1
Male employment rate	-0.004	-0.91	0.000	0.13	0.000	0.09	0.000	-0.01
Male employment rate*Female	0.002	0.55	0.001	0.44	-0.020	-0.68	0.003	1.08
Mother no formal education	0.097	3.26	0.064	3.2	0.078	1.83	0.081	4.48
Mother no formal education*Female	-0.027	-0.66	-0.021	-0.76	0.002	0.06	-0.017	-0.66
Father no formal education	0.041	1.17	0.005	0.29	-0.030	-0.59	0.058	3.61
Father no formal education*Female	-0.028	-0.58	0.003	0.11	-0.059	-1.65	-0.044	-1.86
Mother secondary educ. Or higher	-0.080	-1.56	0.014	0.14	0.024	0.46	-0.126	-3.18
Mother secondary educ. Or higher*Female	0.101	1.41	-0.046	-0.39	0.022	0.69	0.036	0.58
Father secondary educ. Or higher	-0.036	-0.96	-0.036	-1.06	-0.103	-2.45	-0.001	-0.03
Father secondary educ. Or higher*Female	-0.033	-0.63	-0.032	-0.67	0.002	0.91	-0.028	-0.86
Household size	0.006	2.3	0.001	0.47	0.025	0.24	0.005	2.94
Child is a twin	0.092	1.01	0.086	1.18	-0.036	-0.11	0.244	4.09
Older Brothers	0.004	0.34	0.009	1.17	0.029	2.00	0.002	0.22
Older Brothers*Female	-0.022	-1.15	0.003	0.26	-0.027	-1.3	0.000	0.02
Older Sisters	0.004	0.33	-0.001	-0.12	-0.013	-1.11	0.003	0.48
Older Sisters*Female	0.012	0.79	0.011	1.26	0.026	1.56	0.017	1.8
Tests (robust p-values)								
All Female interactions=0	0.0002		0.0000		0.0000		0.0000	
Mother working status equal effect for b/g	0.1357		0.3066		0.7371		0.4550	
Male and Female empl. Rates equal effects b/g	0.4399		0.6496		0.6287		0.4407	
Parental education equal effect for b/g	0.8600		0.8583		0.0526		0.2984	
Sibling composition equal effect for b/g	0.6405		0.3799		0.1310		0.1887	
# observations	6237		15526		5182		14771	
Pseudo R-squared	0.1438		0.1288		0.1388		0.1572	

Source: Author's calculations from NFHS-I and II. Dependent variable: dummy=1 if height-for-age z-score<-2
Does not include Andhra Pradesh, Himachal Pradesh, Madhya Pradesh, Tamil Nadu, and West Bengal.
All regressions also include a constant and state dummies. Standard errors are corrected for clustering.
* See the appendix for a description.

Table 5 - Marginal Effects on the probability of wasting - Probit estimates

	1992-93 (NFHS-I)				1998-99 (NFHS-II)			
	Urban		Rural		Urban		Rural	
	coeff.	t-ratio	coeff.	t-ratio	coeff.	t-ratio	coeff.	t-ratio
Age	0.011	2.78	0.020	9.77	0.005	1.36	0.011	6.03
Age squared	0.000	-2.95	-0.001	-10.08	0.000	-1.78	0.000	-6.05
Asset index (first pr. Comp.)*	-0.028	-4.22	-0.018	-4.24	-0.238	-0.7	-0.014	-3.82
Asset index*Female	0.008	0.98	0.014	2.46	-0.012	-1.56	0.000	-0.08
Muslim	-0.031	-1.4	-0.009	-0.56	-0.006	-0.7	0.009	0.55
Christian	-0.128	-3.01	0.020	0.56	-0.045	-2.17	-0.080	-2.23
Sikh	0.023	0.53	-0.005	-0.17	-0.032	-0.8	-0.027	-0.78
Other Religion (hindu omitted)	-0.040	-0.86	0.002	0.05	-0.067	-1.44	-0.030	-0.6
Mother less than 20 years old	0.000	0.01	-0.006	-0.41	0.021	0.48	-0.001	-0.08
Mother more than 40 years old	-0.075	-1.17	0.036	0.99	0.036	1.09	-0.003	-0.07
Mother does not work	0.084	1.43	-0.005	-0.16	-0.098	-1.03	-0.040	-1.34
Mother does not work*Female	-0.093	-1.18	0.029	0.63	0.059	1.23	0.012	0.3
Mother works away from home	-0.233	-0.8	0.003	0.08	-0.026	-0.39	-0.013	-0.4
Mother works away from home*Female	0.053	0.78	0.016	0.31	0.083	1.39	-0.010	-0.23
Child is a female	-0.061	-0.69	-0.001	-0.74	0.042	0.5	0.002	2.58
Female employment rate	0.001	1.11	0.001	1.43	0.003	2.9	0.000	0.49
Female employment rate*Female	0.000	0.42	0.047	0.29	0.000	-0.01	-0.191	-0.99
Male employment rate	-0.008	-2.24	-0.001	-0.56	-0.007	-1.96	-0.001	-0.7
Male employment rate*Female	0.003	1.01	-0.001	-0.63	0.003	0.77	0.003	1.12
Mother no formal education	-0.029	-0.98	0.014	0.72	-0.010	-0.33	0.027	1.54
Mother no formal education*Female	0.038	0.93	0.022	0.86	0.011	0.25	-0.007	-0.29
Father no formal education	-0.075	-2.54	0.000	0.03	0.038	1.15	0.007	0.45
Father no formal education*Female	0.154	3.26	0.021	0.96	-0.071	-1.79	-0.002	-0.09
Mother secondary educ. Or higher	-0.004	-0.07	-0.142	-1.72	0.036	1.07	-0.012	-0.35
Mather secondary educ. Or higher*Female	-0.032	-0.51	-0.009	-0.08	-0.041	-0.89	0.025	0.46
Father secondary educ. Or higher	-0.038	-0.96	0.088	2.81	-0.048	-1.6	0.002	0.11
Father secondary educ. Or higher*Female	0.017	0.28	-0.116	-2.8	0.007	0.18	-0.073	-2.6
Household size	0.004	2.07	-0.002	-1.86	0.004	1.82	-0.001	-0.93
Child is a twin	-0.005	-0.07	-0.053	-0.92	-0.031	-0.36	0.006	0.12
Older Brothers	0.029	2.29	0.015	2.38	0.015	1.17	0.010	1.59
Older Brothers*Female	-0.023	-1.22	-0.010	-1.14	-0.018	-1.03	-0.012	-1.31
Older Sisters	0.012	1.19	-0.001	-0.09	0.004	0.42	0.008	1.43
Older Sisters*Female	-0.016	-1.05	0.004	0.46	0.017	1.29	-0.003	-0.36
Tests (robust p-values)								
All Female interactions=0	0.0000		0.0079		0.6677		0.0006	
Mother working status equal effect for b/g	0.4686		0.7232		0.4569		0.5850	
Male and Female empl. Rates equal effects b/g	0.5438		0.3345		0.6687		0.3678	
Parental education equal effect for b/g	0.0057		0.0223		0.3817		0.1430	
Sibling composition equal effect for b/g	0.2742		0.4940		0.3160		0.3628	
# observations	6161		15316		5142		14546	
Pseudo R-squared	0.0476		0.0297		0.0477		0.0414	

Source: Author's calculations from NFHS-I and II. Dependent variable: dummy=1 if weight-for-height z-score<-2
Does not include Andhra Pradesh, Himachal Pradesh, Madhya Pradesh, Tamil Nadu, and West Bengal.
All regressions also include a constant and state dummies. Standard errors are corrected for clustering.
* See the appendix for a description.