The Disappearing Nutritional Bias Against Chinese Girls

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Abstract

This paper investigates whether China has reached postnatal nutrient intake equality between boys and girls, despite an exceptionally high ratio of boys to girls at birth, after dramatic technological advances in prenatal sex determination, rapid increases in income, and improved educational opportunities for females. Dominance methods applied to data from the Chinese Health and Nutrition Surveys (selected years 1991 to 2004) reveal no bias in calorie consumption between girls and boys. We find some weak evidence of protein bias toward boys in 1991, but it disappears by 2004.

Keywords: gender bias; undernutrition; lower partial moments; China

JEL classification: I32; O15; N35

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1. Introduction

Modern technology for prenatal diagnostic testing, especially the ultrasound B scan, first became available in China in the 1980s, reached county hospitals by the late 1980s, and rural townships by the mid 1990s (Zhu et al., 2009; Chu, 2001). The new methods dramatically improved the accuracy of prenatal sex determination over older methods, which have been part of traditional Chinese medicine for centuries. Around the same time a strict population control policy was introduced in China, which was seen as crucial to the success of market reforms on which the country had just embarked. The policy comprises a number of laws and administrative rules that regulate family sizes, including marrying and having children at a late age, having easy access to contraception and abortion, and the conditions under which a second child is permitted. The One-Child rule applies mostly to urban residents and state employees while a second child is generally allowed after 5 years, especially if the first child is female (Li, 2004). Coupled with this strict family policy and strong son preferences, prenatal sex selection reached epidemic levels in the late 1980s, as many women opted for abortions upon discovering that the fetus was female. The prevalence of female-selective abortions raised sex ratios at birth (SRB) from 108 boys per 100 girls in 1982 to 121 boys per 100 girls in 2005.

The Chinese government, alarmed by the sharp increase in the SRB, attempted to eliminate prenatal sex selections by banning the use of ultrasound for sex determination apart from medical reasons in 1993, and by enacting the Human Reproductive Technology Ordinance of 2000 that banned sex-selective pregnancy termination. Both efforts appear to have been mostly fruitless, however, because of the rapid privatization and deregulation of health care in China and the widespread use of cheap mobile ultrasound machines during this period.

Several explanations for the changes in the SRB have been offered by other researchers. Banister (2004) notes that while China "is the most extreme example" (p.22), other Asian countries also report abnormal sex ratios. She describes the period preceding the introduction of sex selection technology, the years 1964-77, as the period with "the lowest-ever excess losses of girls in Chinese history (p. 24)" and credits, among other factors, the Communist Party's ideology of gender equality for these low SRBs. Li (2002) also investigates the factors underlying son preference and finds that the One-Child policy, the socioeconomic condition, and the sex ratio of children present are the most influential factors. Zhu et al. (2009), using 2005 survey data, find that "Sex selection abortion accounts for almost all of the excess males (p.1)".[1]

In the current context, with exceptionally high SRBs due to the wide availability of *prenatal* gender screening that strongly favors boys in China, is there still *postnatal* discrimination against Chinese girls? We are not the first to ask whether the recent advances in medical technology that allow parents to make prenatal sex selection decisions could have unintended consequences for the postnatal well-being of girls. Banister (2004) and Echavarri (2007) note that prenatal sex selection should result in fewer unwanted girls, declines in female infanticide, and reductions in "excess" female mortality due to neglect. Our contribution is to investigate a potential form of neglect by testing for inequality across distributions of nutrient intake for girls and boys in China, using micro data from various years between 1991 and 2004.

[1] Some researchers see signs of an improvement in the SRB in the near future; see Das Gupta, et al. (2009) and Zhu, et al. (2009).

Independent of the prenatal sex selection issue, gender bias in development indicators such as nutrient intake has interested economists for several decades. The question here has been whether women and girls, in particular, receive fewer household resources than men and boys in developing countries, which could lead to poorer relative health and nutrition status.[2] Interest in such gender bias is unlikely to diminish, because the millennium development goals (United Nations, 2006) make eradication of extreme poverty and hunger and promotion of gender equality the first and third development goals of the new millennium. Tracking progress on meeting these goals will require methods suited to the detection of movements in the relevant indicators.

The standard approach for testing gender bias in intra-household resource allocation is to measure whether the effect of an additional son on the consumption of adults differs from the effect of additional daughter (Deaton, 1997). Few of these studies have found statistically significant evidence of gender bias, even in data collected prior to the technological advances in sex determination. However, most of these studies have to rely on expenditure data for assignable adult goods that constitute a relatively small share of expenditures.

We take an alternative approach by presenting dominance methods for monitoring gender bias in nutrient intake, using data collected in the China Health and Nutrition Surveys (1991, 1993, 1997, 2000, and 2004). Our data begin just before Deng's economic reforms and continue through a nearly uninterrupted period of expanding economic freedom and rapid growth in China. This growth is reflected in per capita incomes, which grew four-fold from $1479 in 1991 to $5993 in 2004, and could

[2] See Park and Rukumnuaykit (2004) and Dancer et al. (2008) for references to this literature.

have contributed to improved nutrition status for females.[3] In addition, our data cover a period of rapid rural-to-urban migration, with 2004 estimates of rural migrants ranging from 80 to 110 million persons.[4]

We examine whether Chinese females suffer a nutritional disadvantage in calories or protein relative to Chinese males, with particular attention to comparisons of girls and boys. As shown below, we use the lower partial moments (LPM) of nutrition intake distributions, decomposed by subgroup, to make comparisons over the whole range of the distributions and for multiple years. This approach also allows us to pool data from various years and focus on the most vulnerable girls, those from ages zero to two. Finally, to study potential nutrient substitutions, we combine calorie and protein intakes into a simple nutrient index. Our statistical procedure imposes no parametric assumptions about the distribution of intakes among subpopulations or across the whole population. It also allows inferences at any nutrition consumption target values chosen by the researcher.

2. Detecting Gender Bias using Lower Partial Moments

Our approach to comparisons of female and male nutrient intakes owes much to Kakwani (1989), who first applied dominance methods to undernutrition. The particular dominance comparisons that we make are of the LPMs of distributions, decomposed by gender. In the income distribution literature, Deutsch and Silber (1999) and Butler and MacDonald (1987) show that dominance of one set of LPMs over another corresponds to the notion of "economic advantage", or in this application, "nutritional advantage."

[3] See the UN Statistics Division, http://data.un.org/.
[4] Christian Science Monitor, 2004, http://www.csmonitor.com/2004/0123/p08s01-woap.html.

To express the notion of LPMs more formally, let x be a continuous variable with a nutrient probability density $f(x)$, and let $F(x)$ represent the cumulative distribution function (CDF) of X. Let the inverse CDF of X be written $0 \le F^{-1}(p) \le \infty$ and, without loss of generality, let $\tau = F^{-1}(p)$ define target nutrient levels. When p = 0.1, 0.2, …, 1.0, the target nutrient levels become the decile order statistics. Let I_τ^x be an indicator variable such that $I_\tau^x = 1$ if $x \le \tau$ and $I_\tau^x = 0$ otherwise.

Given a target nutrient level τ, we can define the *h*-th partial moment for $x < \tau$ of the density function $f(x)$ as

(1) $$M(\tau;h,X) = \int_0^\tau x^h f(x)dx = \int_0^\infty (x^h I_\tau^x)dF(x) = E\left[(X^h I_\tau^x)\right],$$

where *E* is the expectation operator. Following Butler and McDonald (1987), we define the normalized incomplete moment of X for $x \le \tau$ as

(2) $$\varphi(\tau,h,X) = M(\tau,h,X)/E(X^h),$$

When $h=1$, for instance, the normalized incomplete moment yields the Lorenz ordinates, $\varphi(\tau,1,X) = L(\tau;X)$, the well-known inequality measure. We select fractions of the Recommended Dietary Allowances (RDAs) as our target nutrient levels. Using h = 1 compares the cumulative RDA-equivalent nutrients consumed by all persons below a given fraction of the RDA. The results presented in this paper focus mainly on the h = 0 case, which involves counting the cumulative number of persons below some fraction of the RDA.[5] To check the robustness of these results, we also explore the h = 1 case.

[5] For a more complete discussion of the LPM method and related inference procedures, see Bishop et al. (2003).

The choice between our LPM approach and a typical conditional (regression) approach depends on the question the researcher wishes to address as well as the quality of the data employed. We wish to examine nutrition differences between boys and girls at multiple quantiles of the nutrient distribution, for which the LPM method is well suited. Second, the quality of the nutrient data is superior to that of the potential regressors, particularly rural income. Finally, a conditional analysis leads to a considerable sample size reduction as up to forty percent of the individuals in our data set have missing values for either income or schooling.

3. Application to China

The data for our study come from the China Health and Nutrition Survey (CHNS), conducted by the Carolina Population Center. For our analysis we use the samples from 1991, 1993, 1997, 2000, and 2004.[6] While the CHNS survey was taken in 1989, children were excluded. For the later years, the CHNS is a nationally representative sample from nine regionally-dispersed provinces. The original panel includes 4,400 households with over 16,000 individuals. Our main focus is on children between the ages of zero and 13, with adults ages 18 to 49 as the comparison group.[7] We also provide some pooled results for the most vulnerable group of girls, whose ages are zero to two. In this paper, we use a set of age and gender-specific RDAs sanctioned by the Chinese Nutrition Society (2000). For each specific age and gender group, recommended energy allowances, i.e. calorie intake, represent the *average* needs of individuals. In contrast,

[6] See Liu (2008) for a complete discussion of the CHNS data.
[7] We choose age 13 for our cutoff as the RDAs change frequently between ages 14 and 18. We also considered alternative ages for the cutoffs, with no qualitative change in our findings.

recommended protein allowances are high enough to meet an upper level of requirement variability among individuals within the groups.

Table 1 presents descriptive statistics for our samples. It shows that in 1991 the average girl received 97 percent of her RDA for calories, while the average boy received 96 percent of his RDA, so we find no child gender calorie bias at the means of the calorie samples. Also, in 1991 the average women received 110 percent of her RDA for calories and the average man received 111 percent of his RDA, so we find no adult gender bias at the means of the calorie samples. For protein, boys appear to have some advantage over girls (84 percent of the RDA vs. 79 percent of the RDA in 1991) and men appear to have a slight advantage over women (105 percent vs. 103 percent of RDA in 1991). Shifting to 2004, we again find no evidence of gender discrimination for either children or adults at the calorie consumption means, but we find a significant decline in calorie consumption at the sample means over time, from 110-111 percent in 1991 to 93-94 percent in 2004. This finding is consistent with other studies (e.g., Du et. al., 2004 and Meng et. al., 2004) that report falling nutrition levels over time, usually attributed to rising food prices in China. Indeed, we find declining calorie and protein adequacy for all age groups.

While Table 1 provides interesting insights into gender differences in nutrition, differences in the means may fail to capture important differences in the full distributions. As the previous section showed, constructing the LPMs allows us to uncover any differences in the nutrient distributions for the subgroups and to make corresponding statistical inferences. Table 2 presents the point estimates of the LPMs for calories and protein. As our target nutrient levels, the RDAs are based on the principle that most, if not all, individuals of a population or a specific population group should obtain an adequate nutrient intake to satisfy their daily requirements. We normalize by RDAs to make the comparisons of nutrient intakes across gender and age meaningful. We note that the sample sizes decline our time, from 1484 girls in 1991 to 644 girls in 2004.

In Table 2a we evaluate the RDA-adjusted calorie distribution at seven pre-selected fractions of the RDA for h = 0. The table entries are the proportions of persons below the preset fractions of the RDA. For example, if we set the cutoff at three-quarters of the RDA (0.75) we find that 24.5 percent of the girls and 24.8 percent of the boys lie below this nutritional standard in 1991. To test the hypothesis of equivalent RDA-adjusted distributions of calories, we compare the test statistics at all seven fractions of the RDA.[8] Column 3 provides these test statistics for the girl-boy comparison in 1991. As none of the seven test statistics is greater than 1.96, we can conclude that there is no significant difference in the calorie consumption of boys and girls in 1991. Column 9 reports the corresponding statistics for the 2004 comparison. Again, we find no significant difference in the calorie consumption of boys and girls.

Table 2b presents the RDA-adjusted protein distributions for both children and adults. We find that in 1991 (columns 1 and 2) 49.9 percent of girls and 42.7 percent of boys have protein intakes below the three-quarters of the RDA. The value of the test statistics for this comparison is 2.30, implying that boys enjoy a statistically significant protein advantage over girls in 1991 with respect to this one target level in the

[8] Bishop et al. (2003) present a detailed description of the formal inference tests for the differences of these partial moments over different distributions. Bishop and Formby (1999) provide simulation results for similar types of inference tests. While they recommend using the studentized maximum modulus (SMM) distribution to select critical values (with k= 7, the SMM critical value is approximately 2.5), they note that this can result in a conservative test. We prefer to err in the opposite direction (failing to accept the null of equality when it is "true"), so we use the standard z-table to select our critical values.

distribution. Examining column 9, the 2004 test statistics, we find that none is greater than 1.96. Therefore, we find no evidence that any boy protein advantage in 1991 continues to exist in 2004.

We performed similar dominance tests to those presented in Table 2 for alternative years of the CHNS data (1993, 1997, and 2000). Table 3 summarizes the dominance results for all five years of data. Table 3 uses two types of notation, "NSD," to imply no significant difference at the five percent level and the fraction of the RDA where a significant difference was located. In the first row (1991) reports "NSD" for calories and 0.75 of the RDA for protein; this is simply a summary of the findings of Table 2 for 1991. For all years examined we find no significant differences (NSD) in boy and girl calorie consumption. For protein, boys have an advantage in 1991 and 1997 (at 75 percent of the RDA), and in 2000 (at 95 percent of the RDA). However, it may be the case that girls are substituting calories for protein. To test this hypothesis, Table 3 introduces a "nutrition index," which is a simple average of RDA-normalized calories and protein. Using the nutrition index, we find no significant difference between the nutrient intakes of boys and girls in all five years examined, providing support for the substitution hypothesis.

Some researchers believe that the strongest boy bias should be observed in the first two to three years, because female infanticide – by direct action or by neglect – is rare after age three (Banister, 2004). To test this hypothesis, we pool the five years of CHNS data for ages zero to two. Table 4 provides summary statistics and dominance results from the pooled sample, which includes data from all available years. While baby boys have a higher "Protein/RDA" ratio (0.754 vs. 0.732), baby girls have a higher

"Calories/RDA" ratio (0.816 vs. 0.797). The nutrition index values for baby boys and baby girls, which take both nutrition measures into consideration, are almost exactly the same. Moreover, we find no significant difference in any of the dominance comparisons in Table 4. Overall, these findings suggest that baby boys and baby girls have nearly identical access to nutrition, with some modest substitution of calories for protein among baby girls.

We also tested for the sensitivity of our conclusions to our choices of h = 0 and the preselected targets. Using different targets and performing the comparisons with h = 1 leads to some minor alterations in our findings; it does not alter our overall conclusion of little or no boy bias in nutrition. To explore the possibilities for a conditional analysis, we regressed calories/RDA, protein/RDA and our nutrition index on per capita income, mother's schooling, coastal region, rural residence, female and female*rural variables. The regression results are reported for calories and protein (for years 1991 and 2004) in Table 6. For all years the calorie female coefficients are insignificant. The protein female coefficient is significant for all years except the last year of our sample period (2004). When we replace protein/RDA and calories/RDA with our nutrition index in the conditional analysis, we find no significant differences by gender for all years except 1991. As with our LPM approach, this finding suggests substitution between calories and protein by girls. We emphasize, however, that the conditional analysis suffers from missing data and poor measures of rural income.

4. Discussion

In the previous section we find no evidence that any protein advantage in 1991 continues to exist in 2004. This result is consistent with our contention that prenatal sex selection (as evidenced by exaggerated sex ratios) has improved postnatal outcomes for girls. Nevertheless, the findings presented in this paper are essentially associations. Our case would be strengthened if we could show a larger nutritional bias toward boys in the years prior to the widespread availability of ultrasounds. Unfortunately, nutrient data from this period are scarce. In particular, there are no large-scale representative survey data on Chinese children's nutrition intake status prior to the CHNS with one exception: the 1975 National Growth Survey of Children and Adolescents (NGSCA) (Zhang, 1976). This national survey was conducted by the Capital Institute of Pediatrics and the Ministry of Health and collected ages, heights, and weights for all children on randomly selected residential areas in 9 cities. Using the 1975 NGSCA data and follow-up waves in 1985 and 1995, Li et al. (1999) find that the mean height gap between boys and girls in Beijing, which is one of the cities included in all waves of the survey, is indeed closing for most age groups. For instance, the average height advantages for boys ages 1-3 are 1.4cm, 1.7cm, and 1.4cm respectively in 1975. These advantages drastically declined to 0.6cm, 0.5cm, and 0.6cm in 1995.

Other indirect evidence that suggests the existence of nutritional status bias against girls in China prior to late 1980s. In a comprehensive review of existing survey data on undernutrtion in developing countries, Marcoux (2002) finds that "once noted" anti-female biases in China "are no longer found in national surveys taken at more recent dates" (which refer to late 1980s and early 1990s). Nevertheless, a Chinese Institute of

Nutrition and Food Hygiene 1987 study, which is the earliest one out of four national surveys reviewed by Marcoux (2002), shows that there was a significant sex bias in wasting (low weight for height), stunting (low height-for-age), and underweight (low weight for age) in certain regions of the country.

As argued by Sen (2003), the gender bias in China has been marked by two dramatic changes in two opposite directions during the past decade. On one hand, the female disadvantage in age-specific mortality has declined rapidly. On the other hand, a new gender bias against girls skyrocketed in the fetal stage through gender selective abortions. We can provide some evidence on the changing status of girls in terms of mortality, and when these changes may have come about. Table 5 presents the sex ratios of child mortality for the period 1973 to 2004. The introduction of the strict population control policy in the late 1970's and early 1980's reversed the less-than-one-year mortality rate from a (normal) male disadvantage to a female disadvantage. However, there is some evidence of a turnaround; the female disadvantage decreases between 2000 and 2005. Similarly, the ages 1-4 mortality rate changed from neutral (0.97) in 1973-1975 to a female disadvantage (0.89) in 1981 and back to neutral (1.00) by 1995. Taken as a whole, most researchers (e.g. Das Gupta et al., 2009; Banister, 2004) conclude that, while there is evidence of early life female infanticide, there is no clear evidence that girls have higher mortality rates due to neglect or (relative) lack of access to health care.

There is additional evidence suggesting the improving status of girls in China. For example, primary school enrollment ratios rose from 0.85 in 1980 to 1.0 by 1995. Similarly, Secondary School enrollments rose from 0.70 in 1980 to 0.90 in 1995. Furthermore, Hesketh, Li and Zhu (2005) quote survey results suggesting that attitudes towards girls are changing rapidly in China. In particular, data from the National Family Planning and Reproductive Health Survey shows that 75 percent of respondents from wealthy Jiangsu Province are satisfied with having one child regardless of sex. While responses from poorer provinces were less encouraging, Hesketh et al (2005) believe that gender attitudes are beginning to change in China. Using more recent data, Bi and Ji (2005) examine changes in height and weight of children and adolescents and find no evidence of gender bias in physical growth. Finally, several researchers express optimism that the near future will bring improvements in the SRBs. Das Gupta et al. (2009) note a rapid turnaround in sex ratios in South Korea and find early indications that China may be on the verge of a similar turnaround. Zhu et al. (2009) express optimism regarding prenatal sex selection, citing a slight reduction in the SRB for 2005 as "perhaps indicating the beginning of a reduction in the sex ratios for the future (p. 2)."

5. Conclusions

We argue that an "unintended consequence" of improved sex selection technology in China is reduced postnatal gender discrimination against girls. We use data from the Chinese Household Nutrition Survey to test for gender bias in nutrition in China, focusing particularly on girls and boys (ages 0-13) in selected years from 1991 to 2004. We find no gender bias in calorie consumption for any of the years examined. We find some weak evidence of a bias toward boys in protein consumption in 1991, but we cannot detect any bias in 2004. Moreover, females substitute calories for protein, thus alleviating, in part, the protein bias. For children at the most vulnerable ages (0-2), we find no significant difference in either of calorie or protein distribution by gender. These findings complement other studies that find improvements in post-natal well being for Chinese girls in recent decades, due to a decline in unwanted children. We also recognize that rising incomes and higher levels of education may have contributed to a reduction in the gender bias in nutrition. While there is some hope for future improvement in SRBs, the problem of pre-natal sex discrimination (Sen, 2003) remains severe.

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Table 1. Sample Means for Calories and Protein by Age and Gender

	Girls	***Boys***	***Women***	***Men***
Calories (kcal) in 1991	1685	1781	2548	2970
Calories (kcal) in 2004	1493	1575	2154	2521
Calories/RDA in 1991	0.97	0.96	1.10	1.11
Calories/RDA in 2004	0.82	0.82	0.94	0.93
Protein (g) in 1991	47.9	51.0	71.9	84.0
Protein (g) in 2004	51.5	54.8	63.6	74.8
Protein/RDA in 1991	0.79	0.84	1.03	1.05
Protein/RDA in 2004	0.71	0.76	0.91	0.94

Note: Children are ages 0-13, Adults ages 18-49.

Table 2a. Subgroup Ordinates (h=0) by Gender and Age: Calories, 1991 and 2004

Percent of RDA	1991						2004					
	Girls (1)	**Boys (2)**	**z-score (3)**	**Women (4)**	**Men (5)**	**z-score (6)**	**Girls (7)**	**Boys (8)**	**z-score (9)**	**Women (10)**	**Men (11)**	**z-score (12)**
0.50	0.033	0.038	0.64	0.00	0.00	0.00	0.110	0.139	1.45	0.037	0.039	0.33
0.75	0.245	0.248	0.15	0.082	0.084	0.28	0.449	0.479	0.67	0.267	0.259	-0.56
0.90	0.464	0.465	0.06	0.246	0.248	0.19	0.683	0.673	-0.16	0.500	0.490	-0.44
1.00	0.598	0.602	0.13	0.394	0.419	1.43	0.780	0.786	-0.12	0.639	0.647	-0.31
1.10	0.722	0.730	0.21	0.549	0.578	1.35	0.860	0.858	-0.14	0.765	0.780	0.47
1.25	0.848	0.852	0.08	0.739	0.757	0.70	0.925	0.915	-0.14	0.881	0.891	0.31
1.50	0.941	0.951	0.19	0.907	0.909	0.07	0.980	0.963	-0.22	0.955	0.962	0.21
Sample Size	1484	1680	-	3496	3272	-	644	704	-	2758	2577	-

Note: Boys and girls are ages 0-13, Adults ages 18-49.

Table 2b. Subgroup Ordinates (h=0) by Gender and Age: Protein, 1991 and 2004

Percent of RDA	1991						2004					
	Girls (1)	**Boys (2)**	**z-score (3)**	**Women (4)**	**Men (5)**	**z-score (6)**	**Girls (7)**	**Boys (8)**	**z-score (9)**	**Women (10)**	**Men (11)**	**z-score (12)**
0.50	0.120	0.097	-1.81	0.009	0.009	0.00	0.259	0.251	-0.25	0.073	0.059	-1.80
0.75	0.499	0.427	**-2.30**	0.170	0.156	-1.29	0.624	0.574	-0.90	0.332	0.305	-1.51
0.90	0.711	0.650	-1.50	0.390	0.367	-1.34	0.803	0.726	-1.15	0.555	0.518	-1.47
1.00	0.804	0.764	-0.91	0.543	0.522	-0.99	0.859	0.794	-0.92	0.675	0.648	-0.95
1.10	0.871	0.836	-0.74	0.668	0.650	-0.73	0.904	0.869	-0.47	0.777	0.747	-0.97
1.25	0.930	0.911	-0.38	0.806	0.784	-0.76	0.961	0.928	-0.44	0.865	0.858	-0.22
1.50	0.977	0.974	-0.06	0.922	0.909	-0.40	0.984	0.970	-0.18	0.947	0.940	-0.18
Sample Size	1484	1680	-	3496	3272	-	644	704	-	2758	2577	-

Note: Boys and girls are ages 0-13, Adults ages 18-49. All z-scores greater than 1.96 are in bold.

Table 3: Summary of Dominance Results for Boys and Girls Ages 0-13

	Nutrient Distributions		Nutrient	Sample
Year	**Calories**	**Protein**	**Index**	**Sizes (#boys, #girls)**
1991	NSD	0.75	NSD	1680, 1484
1993	NSD	NSD	NSD	1333, 1504
1997	NSD	0.75	NSD	1128, 1279
2000	NSD	0.95	NSD	999, 1143
2004	NSD	NSD	NSD	644, 704

Note: 0.75 or 0.95 denotes boys dominating girls at these fractions of the RDA. NSD means that there is no significant difference at five percent confidence level. Detailed statistics for 1991 and 2004 are given in Table 2.

Table 4: Summary Statistics and Dominance Results for Ages 0-2 in the Pooled Sample

	Baby Girls	**Baby Boys**
Calories	915	934
Calories/RDA	0.816	0.797
Protein	28.1	28.9
Protein/RDA	0.732	0.754
Nutrition Index	0.774	0.776
Sample Size	374	413
Calorie Dominance?	No Significant Difference ($p < 0.05$)	
Protein Dominance?	No Significant Difference ($p < 0.05$)	
Nutrition Dominance?	No Significant Difference ($p < 0.05$)	

Note: The pooled sample includes data from 1991, 1993, 1997, 2000 and 2004. Babies are defined as 0 – 2 years of age. Raising the top age in this category to 3, 4 or 5 does not alter the dominance results.

Table 5: Sex Ratios of Child Mortality

Years	Mortality < 1 year	Mortality 1-4 years	Mortality < 5 years
1973-75	1.14	0.97	1.06
1981	1.06	0.89	1.00
1990	0.86	0.91	0.87
1995	0.75	1.00	0.79
2000	0.71	0.99	0.72
2005	0.81	0.99	0.81

Source: Das Gupta et al., World Bank Working Paper 4846, 2009

Table 6 – Nutrient Regression Results

Explanatory Variables	**1991**		**2004**	
	Calories	Protein	Calories	Protein
Intercept	0.9578*	0.7839*	0.8603*	0.7225*
	(0.0249)	(0.0234)	(0.0401)	(0.0426)
Per Capita Income/100	0.0021*	0.0046*	0.0006	0.0015*
	(0.0010)	(0.0009)	(0.0004)	(0.0005)
Mother's Schooling	-0.0053*	0.0001	-0.0022	0.0074*
	(0.0018)	(0.0017)	(0.0032)	(0.0034)
Coastal Region	0.0303*	0.0556*	0.0037	0.0712*
	(0.0001)	(0.0001)	(0.0223)	(0.0237)
Rural	0.0142	-0.0134	-0.0243	-0.0762*
	(0.0203)	(0.0191)	(0.0320)	(0.0341)
Rural* Female	0.0497	0.0560*	-0.0479	-0.0464
	(0.0282)	(0.0265)	(0.0426)	(0.0453)
Female	-0.0286*	-0.0874*	0.0261	-0.0162
	(0.0243)	(.02281)	(0.0368)	(0.0392)
N	3045		932	
R-Squared	0.01	0.02	0.01	0.07