INCOME AND CHILD NUTRITIONAL STATUS IN CHINA IN THE 1990s: THREE ESSAYS

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ABSTRACT

Caryn Bredenkamp: Income and Child Nutritional Status in China in the 1990s: Three Essays
(Under the direction of Prof. John S. Akin)

Exploiting the availability of panel data, the first paper examines the trends in child malnutrition in China, both *across* cohorts and *within* cohorts. Descriptive analyses and the results of pooled OLS and probit models provide evidence of a dramatic downward secular trend in underweight, stunting and wasting. While the aggregate picture is overwhelmingly positive, disaggregation of the data by subgroup reveals some disturbing trends: urban-rural, gender and provincial disparities in nutritional status have increased over time; gains appear to have slowed; and, in some provinces, the prevalence of stunting appears to be increasing.

The second paper explores the robustness of the estimated effect of economic status on child nutritional status to alternative income and assets constructs. The internal and external performance of three income measures and eleven asset indices is examined. Then, a series of reduced-form child health demand models, in which these constructs enter the models as income/asset quintiles, is estimated. The analysis reveals that the choice of construct – income or assets – affects results. Further, it provides an indication of the potential direction of bias if one class of measures is chosen over another: income measures tend to produce smaller coefficients than asset measures, but coefficients are more likely to increase progressively, and significantly, with each successive quintile.

The third paper explores the role of income as a determinant of child nutritional status in China, and examines how its effect has been mediated by the one-child policy and changes in the accessibility, cost and quality of healthcare. Pooled OLS and probit models produce large income coefficients. These coefficients shrink when the effects of policies are incorporated into the model, but remain significant. Being an only-child, having shorter traveling times to healthcare facilities and having access to better quality healthcare are all correlated with improved nutritional status. Once community fixed-effects are introduced, however, the effect of income and healthcare becomes insignificant. This pattern is reproduced in pooled OLS and fixed-effects models in which income is instrumented. In addition, it is shown that the one-child policy and healthcare variables are both gender-neutral and income-neutral in their effects on nutritional status.

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PAPER I

TRENDS AND PATTERNS OF CHILD UNDERNUTRITION IN SEVEN PROVINCES OF CHINA, 1991 TO 2000

Between 1987 and 2000, the national prevalence of child undernutrition in China more than halved – the largest decline in that decade in the world. This means that China had effectively attained the Millennium Development Goal (MDG) – of halving the proportion of underweight children under five years by the year 2015 (from its base level in 1990) – before the Millennium Declaration had even been adopted. What these tremendous gains at the national level obscure, however, are significant variations in nutritional improvements at the sub-national level – across provinces and between urban and rural areas – as well as across demographic groups.

Drawing on a panel dataset from China, the China Health and Nutrition Survey, this paper seeks to provide a more nuanced picture of the changes in the nutritional status of Chinese children over the course of the 1990s. With four waves of data (1991, 1993, 1997 and 2000) collected in seven provinces in China, it provides a series of snapshots of the timing and spatial variation of these changes. The panel design enables one to capture changes in nutritional status both *across* cohorts and *within* cohorts. In addition, attention is given to how the magnitude of these changes has varied by demographic subgroup.

1. Conceptualization and measurement of malnutrition

This paper focuses on protein-energy malnutrition (PEM), a condition that is typically described in terms of anthropometric indices, such as underweight, wasting and stunting. While PEM has severe consequences for child morbidity, mortality and mental developmental, it is only one dimension of the malnutrition picture. Child malnutrition in China needs to be understood against the backdrop of trends in other important components of malnutrition, namely micronutrient malnutrition, overweight and obesity.

1.1 Trends in other dimensions of malnutrition: micronutrient malnutrition, overweight and obesity

The prevalence of micronutrient malnutrition, such as Vitamin A deficiency, iron deficiency and iodine deficiency disorders, in China is very low compared to the rest of the developing world. Moreover, the reduction in the prevalence of these deficiencies during the 1990s was very rapid. Already in 1990, the prevalence of Vitamin A deficiency, at 18%, was lower than the average among children under 6 years in the other major developing country regions of the world[1]. By 2000, the prevalence of Vitamin A deficiency had fallen further to 16%[2], with only a small fraction of children suffering from clinical manifestations of the condition such as xerophthalmia (0.26%), night blindness (0.14%) and total clinical Vitamin A deficiency (0.12%) (Tulane University 2004). The total goiter rate among children, a measure of iodine deficiency, fell from 20% to 9% in the short period between 1995 and 1999 (Lin 2000), while over approximately the same period, the percentage of school children with low iodine status (as measured by urine excretion) fell by 75% (Goh 2002) – a phenomenon that is largely attributable to the rapid expansion of the national salt iodization program during the 1990s (Lin 2000). China entered the 1990s with the lowest prevalence of iron deficiency among preschool children of all developing country regions – around 21% – and the prevalence of iron deficiency continued to fall dramatically, halving to 11% by 1995, and dropping even further to 8% by 2000 (UNICEF and Micronutrient Initiative 2004).

Another dimension of the malnutrition picture in China is the increasing prevalence of overweight and obesity. China is described as being in the middle of a "nutrition transition" (Popkin 2004; Popkin and Gordon-Larsen 2004), shifting to a diet, physical activity and body composition profile that is characterized by a growing burden of obesity and nutrition-related non-communicable diseases, especially among females (Popkin

[1] These regions are Sub-Saharan Africa, Middle-East and North Africa, South Asia, South East Asia (excluding China), Central America and Caribbean, Eastern Europe and Central Asia.

[2] There is some evidence that, since 2000, the situation has improved further: in 2004, the Micronutrient Initiative (2004) reported a Vitamin A deficiency prevalence of only 12% among children under 6 years in China.

1998; Monteiro *et al.* 1998). This has been documented in studies of adults (Bell *et al.* 2001), adolescents (Wang *et al.* 1998) and children (Wang *et al.* 2003). It is especially evident among children living in large urban areas and from upper income strata. For example, around 1999, 28% of boys and 14% of girls aged 6 to 12 years in central Beijing (Iwata 2000) and 19% of boys and 11% of girls aged 7 to 12 in the special economic zone of Shenzhen (Guangdong) (Li and Bell 2003) were overweight – prevalence figures that are equivalent to those found in developed countries. The China Health and Nutrition Survey shows that the prevalence of overweight and obesity[3] among children aged 2 to 11 in the seven provinces[4] rose by 44% between 1991 and 2000, from 9% to 13%, increasing at a faster rate among boys (55%) than among girls (33%) and at a faster rate in urban (100%) than rural areas (33%) (own calculations).

1.2 Measurement of protein-energy malnutrition

The anthropometric indicators used in this paper are age- and sex-standardized weight-for-age and height-for-age z-scores[5], constructed on the basis of the 2000 CDC growth charts[6]. Height-for-age z-scores reflect attained growth and, thus, long-term nutritional status. Weight-for-age z-scores are considered a composite measure of the height-for-age and weight-for-height measures, and are expected to capture the consequences of both long-

[3] The prevalence of overweight and obesity was calculated using the body mass index (BMI) cut-offs for children aged 2-18 recommended by the Childhood Obesity Working Group of the International Obesity Taskforce (IOTF) (Cole *et al.* 2000). BMI is an age- and sex-specific weight-for-height index that is calculated as weight(kg)/height(m)^2. The cut-offs for overweight and obesity for children aged 2-18 correspond to BMIs of 25 for overweight and 30 for obesity at age 18. Figures should be treated with some circumspection, however, since there is some evidence that BMI may not be a very good indicator of body fat for children under 6 years (WHO 1995). The correlation between BMI and body fat is also smaller for pubertal children, especially boys (Rodriguez *et al.* 2004). BMI measures were constructed using STATA 8.2.

[4] These seven provinces are Jiangsu, Shandong, Henan, Hubei, Hunan, Guangxi and Guizhou.

[5] Recumbent length-for-age, rather than height-for-age, is used for children under the age of 36 months.

[6] The 2000 CDC growth standards were developed from revisions to the 1977 National Center for Health Statistics (NCHS) growth charts, with the aid of additional national surveys of children in the United States (NCHS 2000). The revisions yielded, *inter alia*, weight-for-age curves that were slightly higher than the 1977 charts for infants under 24 months so that the 2000 growth charts will result in more frequently classifying infants as underweight, and length-for-age curves that were slightly lower than those for the 1977 curves for children older than 6 months resulting in less frequent classification of low length-for-age when using the revised charts. Overall, however, from age 2 to approximately 14 years, the revised weight-for-age and height-for-age percentiles for both boys and girls in 2000 are considered very similar to the 1977 percentiles, but smoother (NCHS 2000).

term and short-term nutritional status (Alderman 2000). For example, a child that is stunted due to chronic malnutrition, but currently receiving optimal nutritional inputs, is still likely to have a low weight-for-age z-score. Likewise, a young child whose long-term nutritional status is good, but who has recently suffered a severe episode of diarrhea or whose household is temporarily food-insecure may also have a low weight-for-age z-score.

Children who have a weight-for-age or height (or recumbent length)-for-age measurement that is less than two standard deviations below the median value of the NCHS/WHO reference group are referred to as underweight or stunted. The terms "severely underweight" and "severely stunted" are used when the measurements are less than three standard deviations below the reference median. 'Mildly underweight" and "mildly stunted" refer to measurements less than one standard deviation below the reference median. For simplicity, this paper may sometimes use the terms "malnourished" or "undernourished" when referring to the concepts of underweight and stunting. A third anthropometric indicator, the weight-for-height z-score, is an indicator of acute short-term undernutrition. Children with weight-for-height z-scores less than -1, -2 or -3 standard deviations below the reference median are referred to as "mildly wasted", "wasted" and "severely wasted", respectively.

Although Chinese infants and children typically fail to achieve the same growth potential as the NCHS reference population, there is sufficient evidence that these standards are appropriate to measure the growth of Chinese children during the 1990s:

First, ethnic Chinese children born in developed countries tend to follow similar growth patterns to the NCHS reference population, suggesting that it is more likely to be environmental rather than genetic factors that contribute to any growth retardation observed among children resident in China. For example, the growth of Chinese Italian children has recently been shown to be within the normal limits of the NCHS reference standard and greater than Chinese children living in China[7] (Toselli 2005), while Chinese Canadian children have been found to achieve lengths equivalent to the NCHS reference median at 9 to 12 months of age[8] (Sit *et al.* 2001).

[7] In fact, weight and length values are *higher* in Chinese Italian children than in Italian children until 12 months of age, and comparable thereafter (Toselli 2005).

[8] The *weight* and *weight-for-height* values of Chinese Canadian children were significantly lower than the median, though. It is hypothesized that feeding practices, rather than genetic potential, contributed to this (Sit *et al.* 2001).

Second, and more generally, children of upper socioeconomic strata in most countries tend to exhibit remarkably similar growth trajectories. An early study (Habicht *et al.* 1974), which admittedly included only one high income group from a *developing* country, found only limited differences across countries in the growth of children under 8 years of age from upper socioeconomic status – not exceeding 3% in height and 6% in weight. More recently, a WHO task group confirmed this finding, showing that, when allowances are made for variations in maternal height and feeding patterns, the growth patterns of infants in seven developed and developing countries had striking similarities[9] (WHO 2000). Regional studies within China, too, highlight some similarities in the growth potential of children from developed countries and children from China's better-off provinces, such as Guangdong (He *et al.* 2001), Beijing (Xu *et al.* 1997) and Hong Kong (Leung *et al.* 2000), but conclude that they still tend to be slightly shorter and lighter than children in the NCHS reference despite have adequate nutrient and energy intake.

Third, the use of the NCHS reference avoids a major disadvantage of relying on country-specific growth standards, namely that any systematic bias or discrimination against specific subgroups is hidden. For example, local standards that are disaggregated by gender and derived from populations in which there is gender bias will lead to the underestimation or obfuscation of gender bias in statistical analysis (Harriss-White 1997).

Fourth, even if the growth potential of Chinese children does differ genetically from that of the children in the United States on whose growth the reference group is based, resulting in a slight overestimation of the true extent of undernutrition among the Chinese, these growth standards remain a useful yardstick with which to measure *changes* in the nutritional status *within* the Chinese population over time – and are thus appropriate for the purposes of this paper.

1.3 The consequences of protein-energy malnutrition

Malnutrition has dire consequences for children's physical and mental development. Isolating the effects of protein and energy deficiencies on health and development outcomes is difficult, confounded by the fact that

[9] The countries included in the study were China, India, Guatemala, Nigeria, Chile, Sweden and Australia. Although the linear growth of Chinese babies did appear to lag slightly behind the other countries, much of that variation was attributed to variations in feeding practices and the rural nature of the Chinese project site – the only rural site in the study.

when food intake is low, the intake of many micronutrients is usually also inadequate (Allen 1994). Nevertheless, it is generally accepted that children who are underweight or stunted are at greater risk of childhood morbidity and mortality, poor physical and mental development, inferior school performance and reduced adult size and capacity for work (WHO 1995).

Malnutrition weakens immune response and aggravates the effects of infection (Pelletier *et al.* 2003) and, so, children who are malnourished tend to have more severe diarrheal episodes and are at a higher risk of pneumonia. Underweight and stunted women are also at more risk of obstetric complications (because of smaller pelvic size) and low birth weight deliveries (ACC/SCN 1997; Agarwal *et al.* 2002). The result is an intergenerational cycle of malnutrition since low birth weight infants tend to attain smaller stature as adults. In addition, malnutrition in early infancy is associated with increased susceptibility to chronic disease in adulthood, including coronary heart disease, diabetes and high blood pressure (Barker *et al.* 2001; Lucas *et al.* 1999; Popkin *et al.* 2001; UNICEF 1998).

Although the precise mechanisms are not clear (Grantham-McGregor and Ani 2001), protein-energy malnutrition in early childhood is also associated with poor cognitive and motor development. The magnitude of the effect is very much dependent on the severity and duration of malnutrition as well as its timing (Black 2003). Moderate protein-energy malnutrition of long-term duration has worse consequences for cognitive development than transient severe undernutrition. With respect to timing, it is nutritional status in the period between the last trimester of pregnancy and two to three years of age that is most important for mental development.

Owing to its effects on physical, cognitive and mental development, undernutrition limits educational attainment – especially in areas where undernutrition is particularly severe or widespread. Malnutrition has been shown to be associated with a reduced capacity to learn (as a result of early cognitive deficits or lowered current attention spans), fewer total years of schooling (Alderman *et al.* 2005), a reduced probability of ever attending school, particularly for girls (Alderman *et al.* 2001), later entry into school, more repeated grades (Glewwe *et al.* 2001), delays in school completion (Alderman *et al.* 2003), and lower overall high school completion rates (Daniels and Adair 2004).

2. The CHNS data

Data are drawn from a large-scale household survey conducted in nine provinces of eastern China from 1989 to 2000, the China Health and Nutritional Survey (CHNS). This is a panel dataset with five waves of data, from 1989, 1991, 1993, 1997 and 2000.

2.1 Sampling

A multistage, random cluster process was used to draw the sample in each of the provinces. Counties in each province were stratified by income (low, middle, and high) and a weighted sampling scheme was used to randomly select 4 counties in each province. In addition, the provincial capital and a lower income city were selected when feasible. Villages and townships within the counties, and urban and suburban neighborhoods within the cities, were selected randomly.

The first CHNS wave, in 1989, surveyed 15,917 individuals. In CHNS 1991, individuals belonging to the original sample were re-interviewed resulting in a sample of 14,778 individuals. In CHNS 1993, all the new households formed from the 1991 sample who resided in sample areas were added, resulting in a total of 13,893 individuals. In CHNS 1997, all newly-formed households who resided in sample areas were once again included, and additional households and communities were selected to replace those no longer participating. Also, Heilongjiang province replaced Liaoning province, bringing the sample size to 14,426 individuals in 1997. In CHNS 2000, newly-formed households, replacement households, and replacement communities were again added, and Liaoning province returned to the study. In 2000, the sample consisted of 15,648 individuals. The number of primary sampling units (i.e. communities) increased from 190 in 1991 and 1993 to 216 in 2000.

This particular analysis will cover the years 1991 through 2000 because in 1989 anthropometric data were not collected for children over the age of six years. It will cover the seven provinces of Jiangsu, Shandong, Henan, Hubei, Hunan, Guangxi and Guizhou. The provinces of Liaoning and Heilongjiang are excluded because the former lacks data for 1997 and the latter lacks data for 1991 and 1993.

2.2 Data collection

Anthropometric data were collected by trained nutritionists, who were experienced in the collection of national survey data. They had also received additional training prior to CHNS data collection to enhance the degree of precision and level of accuracy of their measurements. In the field, standardized equipment provided for the purposes of the survey was used and anthropometric measurements checked intermittently by field supervisors (CAPM 1993). Height data were recorded to the closest tenth of a centimeter using a horizontal height measurement bed or portable board for children under the age of three and a height scale for older respondents. Weight data were recorded using a physician's scale and recorded in kilograms to two decimal places.

2.3 Data cleaning

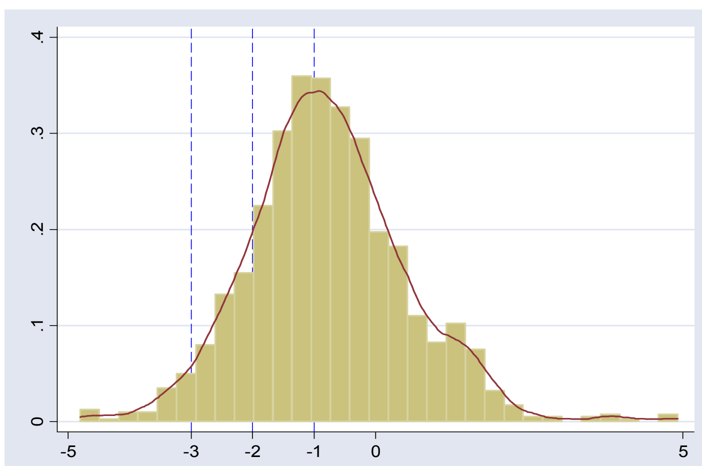
The analysis in this paper utilizes the new CHNS Longitudinal Master Files that have been subjected to extensive cleaning to enhance the accuracy of anthropometric data. Cleaning included the correction of birth date and gender variables, as well as the correction of weight and height variables, based on comparisons of the recorded values with CDC growth charts and with each child's own recorded observations in prior and subsequent waves of the CHNS. Adhering to the standard fixed exclusion ranges for implausible weight-for-age and height-for-age z-scores recommended by the WHO (1995), weight-for-age z-scores that are less than -5 and greater than +5, and height-for-age z-scores that are less than -5 and greater than +3 were excluded prior to analysis.

After cleaning and pooling of data from seven provinces across four years, there are more than 7,500 anthropometric observations (specifically, 7,644 weight-for-age data points and 7,588 height-for-age data points) for children under the age of twelve (under 144 months). The sample contains 2,457 children in 1991, 2,173 children in 1993, 1,739 children in 1997 and 1,275 in 2000. A decrease in sample size of this magnitude is to be expected in a panel study in light of the sharp decline in crude birth rates in China over this period (UNESCAP 2006a)[10]. More details on sample size and characteristics are in Table 5 in the Appendix.

3. Malnutrition in China today – patterns and recent trends

The profile of protein-energy malnutrition in China today is one where the distributions of children's age- and sex-standardized weight and height remain firmly to the left of the global reference standard. In 2000, in the seven provinces included in this study, 15% of children under the age of 12 were underweight and 21% were stunted. Severe underweight afflicted 3% of all children, while 5% suffered from severe stunting. Mild underweight and stunting, which also has important consequences for health, affected almost half of all children, at a prevalence of 45% and 50% respectively.

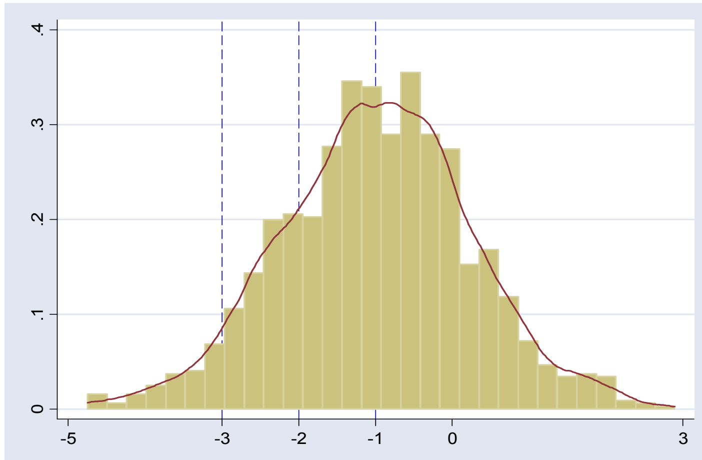
Figure 1: Distribution of weight-for-age z-scores of children under 12, 2000



Source: China Health and Nutrition Survey, 2000
Note: Sample size is 1,275; data are from seven states; dashed lines at -1, -2 and -3 standard deviations represent mild underweight, underweight and severe underweight cut-offs, respectively; the kernel-density plot is computed using an Epanechnikov kernel.

[10] Across all of China, the crude birth rate fell from 21 per 1,000 in 1990 to 13.8 per 1,000 in 2000 (China Statistical Yearbooks 1991-2001; UNESCAP 2006a; UNICEF 2006). In some provinces the decline was even sharper. In Hubei province, for example, the crude birth rate fell by just over 60%, from 24.7 to 9.71, between 1990 and 2000 (UNESCAP 2006b).

Figure 2: Distribution of height-for-age (or recumbent length-for-age) z-scores of children under 12, 2000



Source: China Health and Nutrition Survey, 2000
Note: Sample size is 1,258; data are from seven states; recumbent length-for-age is used for children under 36 mths; dashed lines at -1, -2 and -3 standard deviations represent mild underweight, underweight and severe underweight cut-offs, respectively; the kernel-density plot is computed using an Epanechnikov kernel.

Whether measured in terms of underweight or stunting, it is clear that the nutritional status of Chinese children is far superior to that of other countries in the East Asia and Pacific region (see Table 1), and certainly better than in South Asia and Sub-Saharan Africa. In fact, their nutritional status most closely resembles children in the Latin American and Caribbean region, but with a slightly lower prevalence of stunting and a slightly higher prevalence of underweight.

Nutritional status in the seven CHNS provinces is poorer than in China as a whole. In the under-five age group in 2000, the prevalence of underweight was 14% (compared to 10% for all China) and the prevalence of stunting was 23% (compared to 23% for all China). These figures are very similar to those recorded for the Meddle East and North Africa region.

Table 1: International comparison of the prevalence of undernutrition among children under 5

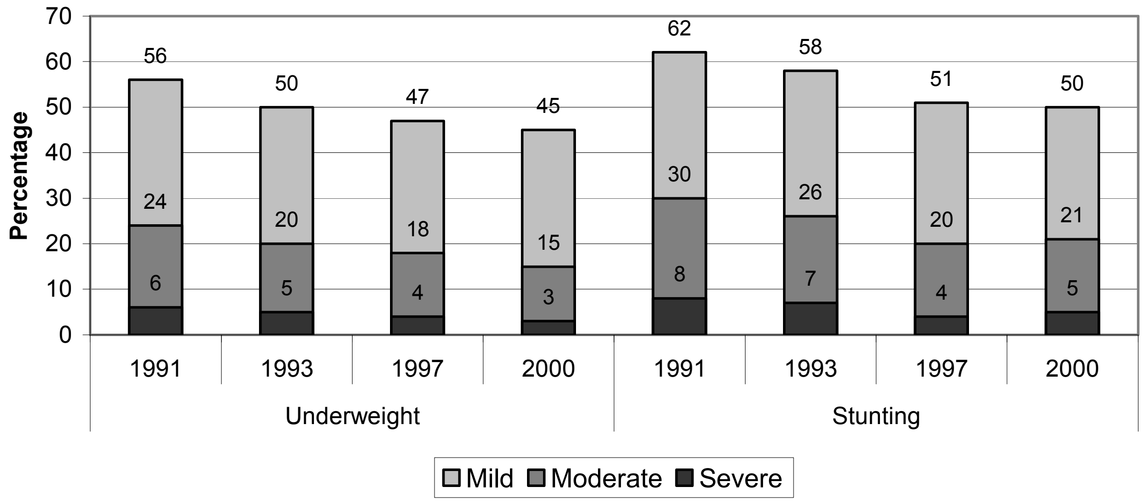
	% of under-fives (1995-2001) suffering from	
Region	Underweight	Stunting
World	27	32
Latin American and Caribbean	8	16
Middle East and North Africa	14	22
Sub-Saharan Africa	29	40
South Asia	46	45
East Asia and Pacific (incl. China)	17	21
China	10	14
CHNS provinces	14	23

Source: UNICEF 2003

Note: Data refer to the latest available information between 1995 and 2001 and so may not be strictly comparable across regions. Data on China are from 2000, from the WHO Global Database on Child Growth (2004), and cover all provinces of China.

The most striking element of the malnutrition picture in China is the rapidity with which the prevalence of malnutrition has fallen in recent years. In the seven CHNS provinces, among children under 12, the prevalence of moderate underweight fell by 37.5%, from 24% in 1991 to 15% in 2000; the prevalence of moderate stunting fell by 30%, from 30% to 21%. The prevalence of severe malnutrition approximately halved, with the prevalence of severe underweight falling from 6% to 3% and the prevalence of severe stunting falling from 8% to 5%.

Figure 3: Trends in the prevalence of underweight and stunting, children under 12, 1991-2000



Source: China Health and Nutrition Survey, 1991-2000
Note: The prevalence of undernutrition in 2000 is statistically different from that in 1991, regardless of the indicator of undernutrition used

The magnitude of these reductions, measured for the seven CHNS provinces, is much smaller than those that have been observed in national surveys of China. WHO data for children *under five* show that between 1992 and 2000 the prevalence of underweight fell from 17% to 10% and the prevalence of stunting fell from 31% to 14% (WHO Global Database on Child Growth 2004; original data for 1992 from Ge 1995, for 1998 from CAPM and SSB 2000, and for 2000 from CAPM and SSB 2001; see also De Onis *et al.* 1993). This implies a percentage reduction of 41% in underweight and 56% in stunting, compared to reductions of 26% and 18% among children of a similar age (under five) in the CHNS provinces over a similar time period[11].

[11] Among children *under five* included in the CHNS, the prevalence of underweight fell from 19% to 14% and the prevalence of stunting fell from 28% to 23% between 1991 and 2000.

4. Demographic differentials in the current prevalence of and recent trends in protein-energy malnutrition

Focusing on the tremendous improvement in aggregate nutritional status obscures changes in the relative position of particular subgroups. During the nine years under analysis, there were significant variations in the magnitude of the changes in the prevalence of undernutrition at the sub-national level – across provinces and between urban and rural areas – as well as across demographic subgroups.

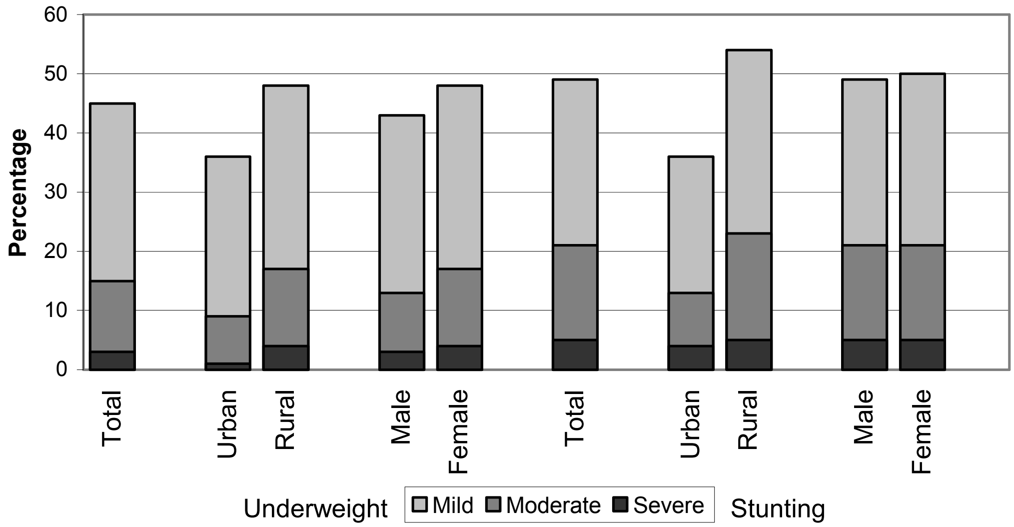
4.1 Current prevalence, by demographic subgroup

Disaggregation of the 2000 CHNS data by urban-rural location, gender and age shows that there are particular demographic groups that are more likely to be undernourished than others.

Urban-rural location: Undernutrition is much more prevalent in rural areas, where 17% of children are underweight and 24% are stunted, than in urban areas, where the figures are 9% and 13% respectively.

Gender: There are also statistically different gender differences in the prevalence of underweight - 17% of girls are underweight compared to only 13% of boys. These gender differences do not appear to manifest themselves in height differentials, however.

Figure 4: Disparities in the prevalence of undernutrition among children under 12, by socioeconomic subgroup, in 2000



Source: China Health and Nutrition Survey, 2000
Note: All urban-rural differences are statistically significant at the 5% level except for severe stunting; male-female differences are only statistically significant for moderate underweight. Exact prevalence figures are in Table 6 in the Appendix

Age: The age-wise pattern of underweight and stunting is another important dimension of undernutrition in China. Weight-for-age z-scores decline steadily, although at a decreasing rate, as children age. While this may, in part, reflect a secular trend of improvements in nutritional status over time, this pattern is in line with growth patterns observed, at least in the first few years of life, among children in many other parts of the world: growth retardation tends to originate early in life, and growth continues to falter as children age, with much of this early damage being irreversible (ACC/SCN 2004; Shrimpton *et al.* 2001). Moreover, this pattern is observed for every cohort of children in the CHNS sample (see Figure 6 in section 4.2), suggesting that there is an age trend that is independent of the secular trend. A similar, but somewhat more erratic, trend is observed for height-for-age z-scores.

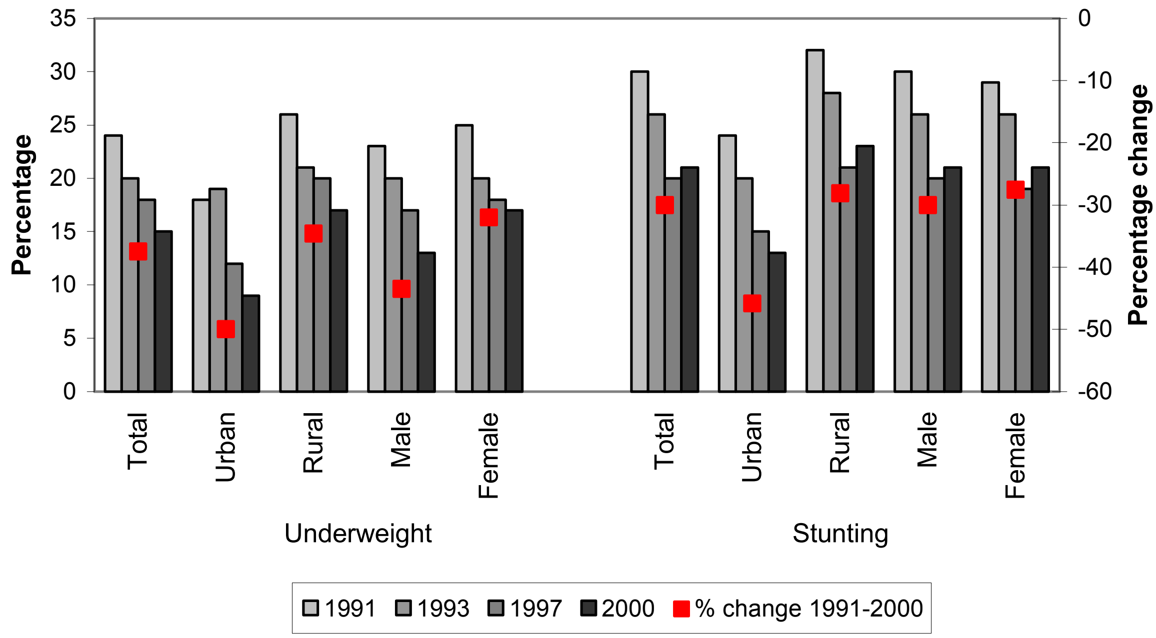
4.2.Recent trends in the prevalence of protein-energy malnutrition, by demographic subgroup

While the overall trend of a declining prevalence of malnutrition characterized all subgroups of the population, nutritional gains appear to have accrued disproportionately to particular subgroups. Although there are differences in magnitude, the underweight and stunting indicators reveal a similar pattern of change.

Urban-rural location: Prevalence declined at a greater rate in urban than in rural areas so that the decadal gains were regressive in nature. The prevalence of (moderate) underweight fell by 50% in urban areas between 1991 and 2000, compared to 35% in rural areas. Similarly, the prevalence of (moderate) stunting fell more dramatically in urban areas (46%) than in rural areas (25%).

Gender: The gender differentials in underweight widened over the course of the 1990s. The prevalence of (moderate) underweight among boys fell by 43% between 1991 and 2000, while the prevalence among girls fell by only 32%. There was no marked change in gender differentials in the prevalence of stunting, however.

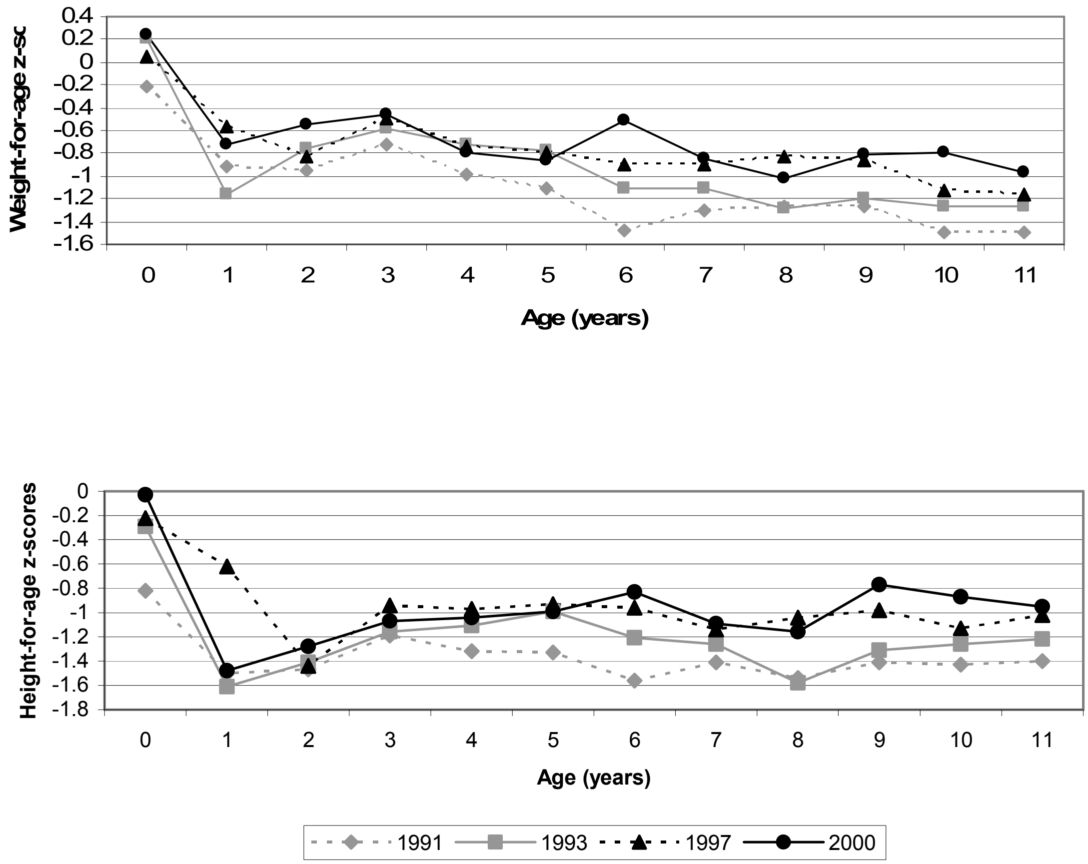
Figure 5: Prevalence, and percentage change in prevalence, of underweight and stunting, by subgroup, 1991-2000



Source: China Health and Nutrition Survey, 1991-2000
Note: In each subgroup, the prevalence of underweight and stunting in 2000 is significantly different from in 1991. Exact prevalence figures are in Table 7 in the Appendix

Age: The trend in weight-for-age and height-for-age z-scores is, by and large, robust to disaggregation by age. The figures below clearly show that the growth trajectories of children under 12 have shifted up over time, suggesting improvements in nutritional status at all ages. It also appears that the rate at which weight-for-age z-scores decline as children age has become less pronounced over time. While it is difficult to ascertain whether specific age groups have benefited relatively more over time, partly because of the bumpier curves associated with smaller sample sizes, the gains to the very youngest children do appear to be relatively smaller than in other age groups.

Figure 6: Trends in weight-for-age and height-for-age z-scores, by age, 1991-2000



Source: China Health and Nutrition Survey, 1991-2000

5. Provincial variation in protein-energy malnutrition

There is great variation in the patterns and trends in malnutrition across the provinces of the country, and even across the seven provinces in the CHNS sample. In 2000, by the WHO classification[12], the severity of malnutrition ranged from unambiguously low in Jiangsu to unambiguously high in Guizhou. In other provinces,

[12] The WHO classification (WHO 1995) is used to classify the severity of malnutrition in a population as "low", medium", "high" or "very high", based on the percentage of the population under the age of five that is (moderately) underweight, stunted or wasted. For underweight, a prevalence range of less than 10% is considered low severity, 10%-19% is medium severity, 20%-29% is high and more than 30% is very high; for stunting, a prevalence range of less than 20% is considered low severity, 20%-29% is medium severity, 30%-39% is high and more than 40% is very high; for wasting, a prevalence range of less than 5% is considered low severity, 5%-9% is medium severity, 10%-14% is high and more than 15% is very high.

the severity of undernutrition when measured by the stunting indicator was low, on average, but medium when measured by the underweight indicator. It is only in Shandong where the prevalence of stunting fell into a higher category of classification than the prevalence of underweight.

Table 2: Classification of provinces by severity of malnutrition, 2000

	Underweight			**Stunting**		
	Low <10%	**Medium 10%-19%**	**High 20%-29%**	**Low <20%**	**Medium 20%-29%**	**High 30%-39%**
Jiangsu	5%			7%		
Shandong	2%				20%	
Henan		10%		18%		
Hubei			24%	17%		
Hunan		11%		18%		
Guangxi		13%		18%		
Guizhou			28%			40%

Source: China Health and Nutrition Survey, 2000

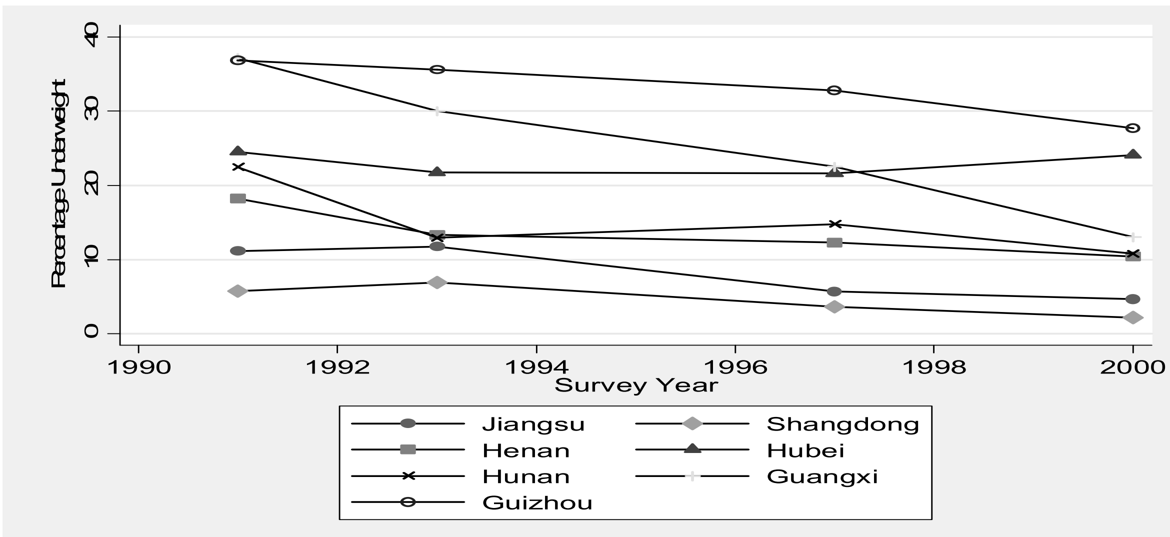
An analysis of underweight and stunting data in these provinces across the four years yields some interesting results:

The first point of note is the great variation in the magnitude of the pace of change across provinces, even within the limited geographic area of the CHNS sample (see Figure 7 and Figure 8 below). The neighboring provinces of Guangxi and Guizhou, for example, both had a similar prevalence of underweight in 1991 (37%) but within the following ten years, this prevalence had fallen by 65% in Guangxi, but only by 24% in Guizhou. A similar picture emerges from the stunting indicators, with the prevalence of stunting falling by 47% and 13%, respectively, in the two provinces.

Second, despite the aggregate downward trend in the prevalence of stunting, in three of the seven provinces (Guizhou, Shandong and Henan) the prevalence of stunting appears to have *increased* between the last two survey years.

Third, there is evidence of growing health disparities across provinces and some of the improvements in nutritional status have been regressive. For example, the two provinces that had the lowest prevalence of underweight in 1991, namely Jiangsu and Shandong, appear to have experienced two of the largest percentage reductions in the prevalence of underweight, so that the alleviation of malnutrition was greatest in the provinces that were already fairly advantaged to begin with[13].

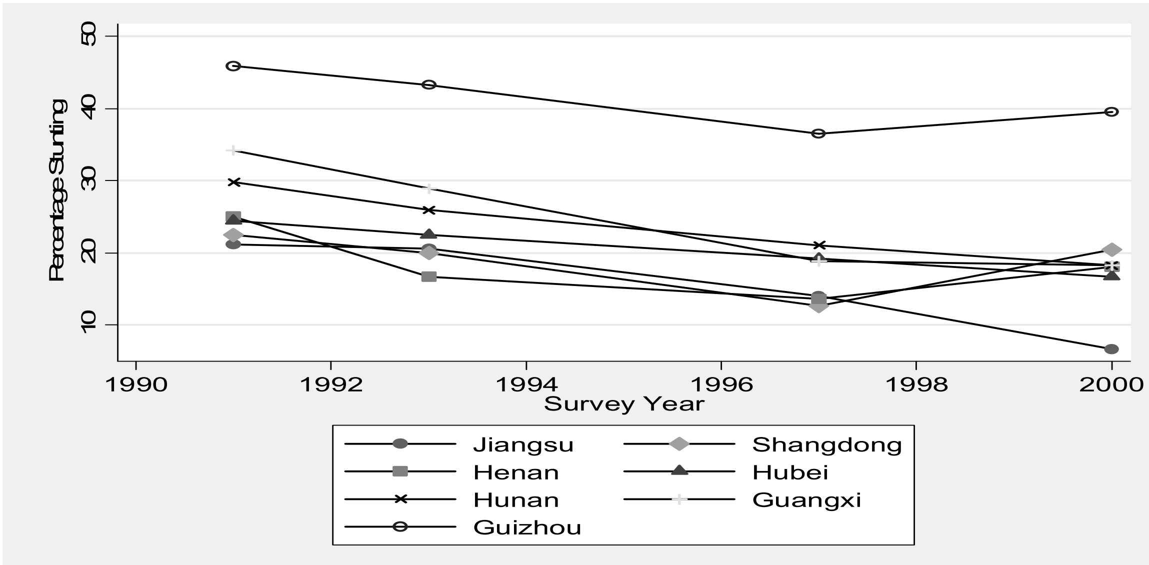
Figure 7: Trends in the prevalence of underweight, by province, 1991-2000



Source: China Health and Nutrition Survey, 1991-2000
Note: Exact prevalence figures are in Table 8 in the Appendix

[13] Typically, the opposite holds: percentage reductions in the prevalence of malnutrition tend to be greatest in those areas with an initial high prevalence.

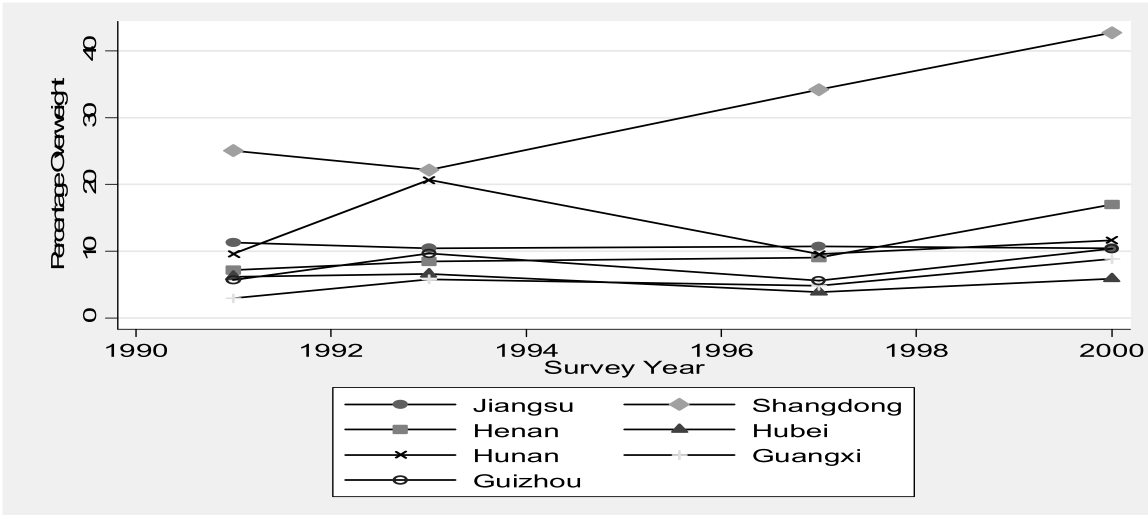
Figure 8: Trends in the prevalence of stunting, by province, 1991-2000



Source: China Health and Nutrition Survey, 1991-2000
Note: Exact prevalence figures are in Table 8 in the Appendix

Fourth, it can be seen that the prevalence of stunting is typically much higher than the prevalence of underweight, and as time has passed, the gap between the two appears to be increasing. In other words, the prevalence of underweight has fallen much faster than the prevalence of stunting. Moreover, there is substantial inter-provincial heterogeneity in the extent of this effect. In Jiangsu, for example, the prevalence of stunting is only slightly higher than the prevalence of underweight, but in Shandong the prevalence of stunting is about ten times as high as the prevalence of underweight. One likely explanation for this phenomenon is the change in dietary quality, and specifically, the shift from a high-carbohydrate to a high-fat, energy-dense diet that has been observed over this period. Indeed, in the CHNS provinces, from 1989 to 1997, kilocalorie intake fell and edible oil intake increased by 76% (Du *et al.* 2004). Children who have low weight- and height-for-age early in life and follow this type of diet are unlikely to reach their linear growth potential, but may still gain substantial weight. BMI data support this hypothesis, highlighting the disproportionate gain in weight relative to height during the 1990s: mean BMI-for-age z-scores for children aged 2 to 11 increased by 29% from 1991 to 2000, which is equivalent to a 44% increase in the prevalence of overweight (own calculations). The graph below shows these changes at the provincial level, including the dramatic increase in overweight in Shandong.

Figure 9: Trends in the prevalence of overweight among children aged 2 to 11, by province, 1991-2000



Source: China Health and Nutrition Survey, 1991-2000
Note: Overweight figures are calculated using the standard BMI categorization, and in this graph includes those children categorized as both overweight and obese. BMI categories are only valid from the age of two

6. Persistence of undernutrition within cohorts across time

The panel structure of the CHNS enables one to explore the extent to which undernutrition persists within a cohort of children over time. Tracking, defined as the maintenance of a certain status (such as underweight or stunting) or a relative position within a distribution of values in a population over time (Twisk *et al.* 1994; Wang *et al.* 2000), provides an indication of the changes in the nutritional status of a particular group of children over a period. One way to observe the tracking phenomenon is to calculate the percentage of children who were undernourished in a previous period that remain undernourished in a subsequent period[14].

In the seven provinces of the CHNS, the extent of tracking of underweight and stunting is not very large, suggesting that over the 1991 to 2000 period a large percentage of children "moved out" of a state of undernutrition. For about half of the children in this sample, undernutrition appears to be a transient state.

For example, of the 422 children who were underweight in 1991 and for whom there were anthropometric data in 1993, 60% were still underweight in 1993; of the 177 children who were underweight in 1991 and for

[14] Other methods include the calculation of the correlation coefficient between a measurement at an early age and one later in life, the use of multivariate logistic regression, and a generalized estimating equations (GEE) approach (Twisk *et al.* 1997)

whom there were anthropometric data in 1997, 54% were still underweight in 1997; and, of the 67 children who were underweight in 1993 and for whom there were anthropometric data in 2000, 51% were still underweight in 2000. This suggests that, at least for some children in the sample, underweight is far from irreversible: a fairly low percentage of children remained undernourished in consecutive periods and the longer the duration of the period under analysis, the greater the percentage of children that managed to transition out of malnutrition[15].

A similar pattern is observed for the stunting indicator, but with greater evidence of tracking between 1991 and 1993. Also, the extent of tracking of underweight and stunting appears to have been much greater between 1997 and 2000 (64% and 62% respectively) than between 1993 and 1997 (60% and 55% respectively) and between 1991 and 1993 (60% and 51% respectively).

Disaggregation by demographic subgroup yields information about which groups are more at risk of persistent undernutrition.

Gender: Throughout the 1990s, a greater percentage of girls suffered from persistent underweight than boys. However, it was the boys who tended to exhibit more persistent stunting – except towards the close of the decade when a greater percentage of boys than girls managed to recover their growth trajectories.

Urban-rural: A rather ambiguous pattern is found when the data are disaggregated by urban and rural areas. From 1991 to 1993, a greater percentage of urban than rural children suffered from persistent underweight, but a greater percentage of rural than urban children suffered from persistent stunting. From 1993 to 1997, the opposite is true: a slightly greater percentage of rural children suffered from persistent underweight, but a greater percentage of urban than rural children suffered from persistent stunting. From 1997 to 2000, the picture is unequivocal: both persistent underweight and persistent stunting were much more prevalent among rural than urban children.

[15] Some caution needs to be used in interpreting the potential for growth recovery suggested by these figures, especially as the period under analysis increases. Since the potential for growth recovery is greatest when children are very young, e.g. under three, we should expect to see a smaller percentage of persistent undernutrition for children initially observed at younger ages. Also, as the period under analysis increases, the prevalence of persistent undernutrition should fall. For example, children who were observed in both 1991 and 2000 would have been between under the age of two years in 1991 and, since the probability of growth recovery declines as children age, we find that only 48% of them remain stunted over that period. In summary, the higher the initial age at which children were observed and the shorter the time interval over which they were observed, the less likely we are to find evidence of recovery.

Table 3: Tracking of underweight and stunting over time, children under 12, 1991 - 2000

	Underweight			Stunted		
	1993	1997	2000	1993	1997	2000
Of children observed to be underweight/stunted in 1991						
All	60% (n=422)	54% (n=177)	51% (n=67)	68% (n=535)	51% (n=246)	48% (n=83)
Boys	50% (n=219)	53% (n=90)	45% (n=31)	69% (n=277)	57% (n=134)	42% (n=43)
Girls	62% (n=203)	55% (n=87)	56% (n=36)	66% (n=258)	45% (n=112)	55% (n=40)
Urban	63% (n=83)	39% (n=33)	25% (n=12)	62% (n=110)	49% (n=43)	31% (n=13)
Rural	60% (n=339)	58% (n=144)	56% (n=55)	69% (n=425)	52% (n=203)	51% (n=70)
Of children observed to be underweight/stunted in 1993						
All		60% (n=178)	52% (n=69)		55% (n=267)	53% (n=109)
Boys		62% (n=99)	49% (n=37)		55% (n=150)	49% (n=57)
Girls		58% (n=79)	56% (n=32)		54% (n=117)	58% (n=52)
Urban		45% (n=38)	17% (n=12)		63% (n=43)	60% (n=15)
Rural		64% (n=140)	60% (n=57)		53% (n=224)	52% (n=94)
Of children observed to be underweight/stunted in 1997						
All			64% (n=119)			62% (n=166)
Boys			59% (n=63)			61% (n=89)
Girls			70% (n=56)			64% (n=77)
Urban			56% (n=18)			50% (n=38)
Rural			65% (n=101)			66% (n=128)

Source: China Health and Nutrition Survey, 1991-2000

Note: The interpretation of the table is as follows: in the three left-hand cells of the top row, for example, we see that of the 422 children who were underweight in 1991 and for whom there are anthropometric data for 1993, 60% were still underweight in 1993; similarly, of the 177 children who were underweight in 1991 and for whom there are anthropometric data for 1997, 54% were still underweight in 1997; and, of the 67 children who were underweight in 1991 and for whom there are anthropometric data for 2000, 51% were still underweight in 2000. The sample size changes because calculations are restricted to those children who were underweight in the first period and also have observations in a subsequent period.

7. Isolating secular trends in child nutritional status

The analyses in the previous sections do not allow one to isolate how much of the change in standardized linear growth or weight over time is due to children's aging and how much is due to a secular trend in nutritional status – a trend that could arise from changes in other factors associated with children's nutritional status, such as diet quality, caloric intake, disease incidence, household income growth and fertility.

The use of a linear regression model that pools the different waves of data and includes both the age variable and year dummies enables one to separate these effects from each other. The year dummies also effectively control for changes in the sample distribution across different time periods – i.e. for the possibility that the level of urbanization, the gender ratio and the age composition of the sample may have changed. In addition, the inclusion of provincial variables allows for spatial variation in these potentially confounding factors. The province of Shandong is the omitted category since, as has been shown in the descriptive analyses, it has some of the better nutritional indicators in the sample. The specification of the model, without interaction terms, assumes that the effect of each explanatory variable on the z-scores has remained constant over time. Huber standard errors are used to correct for heteroskedasticity and clustering at the community level.

The pooled OLS model (see Table 4) confirms the clear upward secular trend in nutritional status over time. Controlling for age, gender, urban-rural location and provincial variation, height-for-age z-scores increased by 0.17 standard deviations between 1991 and 1993; by 0.38 standard deviations between 1991 and 1997; and, by 0.45 standard deviations between 1991 and 2000. Weight-for-age z-scores increased by 0.19 standard deviations between 1991 and 1993; by 0.36 standard deviations between 1991 and 1997; and, by 0.5 standard deviations between 1991 and 2000[16]. Individually, each year dummy is statistically significant at the 1% level and the coefficients increase by fairly large orders of magnitude for each subsequent wave of the survey. According to the Wald test, the difference between the 1997 and 2000 coefficients in the height-for-age z-score model is not statistically significant, though, suggesting that the significant gains to linear growth occurred in the earlier part of the decade.

The secular improvements in nutritional status are also evident in pooled probit models that employ the policy-relevant cut-points of stunting and underweight as dependent variables (see Table 9 in the Appendix). The probability of stunting fell by 4 percentage points between 1991 and 1993 and by 9 percentage points between 1991 and 1997, after which it ceased to decline significantly[17]. The probability of underweight fell progressively throughout the decade: by 4 percentage points between 1991 and 1993, by 7 percentage points between 1991 and 1997, and by 10 percentage points between 1991 and 2000.

The mean BMI-for-age z-score in the pooled OLS model is 0.256 of a standard deviation larger in 2000 than in 1991. This suggests that in addition to being, on average, half a standard deviation taller at the end of the decade than at the start, Chinese children had also gained considerable body mass.

[16] The magnitudes of the reduction in height-for-age z-scores and the reduction in weight-for-age z-scores over the course of the decade were similar – equivalent to about half a standard deviation. Thus, although these two anthropometric measures capture different physical attributes of children, it appears that in their role as indicators of nutritional status in China both point to similar conclusions about the trends in child nutritional status over time.

[17] Chi-squared tests of the equality of the 1997 and 2000 coefficients in the stunting model show that they are not statistically significant.

With respect to other covariates included in the specification, the OLS model confirms the trends and patterns discussed elsewhere in this paper. As children age, their nutritional status declines, but at a decreasing rate. Despite the fact that z-scores are already sex-standardized, boys weight-for-age z-scores are significantly higher than that of girls, although their height-for-age z-scores are not. This suggests that boys in the CHNS provinces are, on average, heavier than girls – a result that is confirmed by the significant male coefficient in the BMI-for-age regression. Very large statistically significant differences are observed between children living in urban and rural areas, and the nutritional status of the former is approximately 0.4 standard deviations better than the latter. Provincial disparities in malnutrition are enormous: differences in height-for-age z-scores vary by more than one standard deviation and differences in height-for-age z-scores vary by more than one-and-a-half standard deviations between the best-off (Shandong) and the worst-off (Guizhou) of the seven provinces in this sample.

Table 4: Pooled OLS model of the correlates of height-for-age, weight-for-age and BMI-for-age z-scores

	Height-for-age	Weight-for-age	BMI-for-age
Age (mths)	-0.101	-0.155	-0.199
	(0.022)***	(0.022)***	(0.031)***
Age-squared	0.006	0.006	0.006
	(0.002)***	(0.001)***	(0.002)***
Male	0.029	0.113	0.124
	-0.037	(0.030)***	(0.030)***
Urban	0.407	0.398	0.176
	(0.080)***	(0.069)***	(0.050)***
Year dummies			
1993	0.167	0.193	0.109
	(0.026)***	(0.038)***	(0.046)**
1997	0.38	0.362	0.167
	(0.040)***	(0.047)***	(0.051)***
2000	0.445	0.497	0.256
	(0.049)***	(0.052)***	(0.057)***
Province dummies			
Shandong (omitted)			
Jiangsu	-0.187	-0.589	-0.51
	-0.174	(0.101)***	(0.157)***
Henan	-0.32	-0.729	-0.632
	(0.171)*	(0.120)***	(0.149)***
Hubei	-0.507	-1.18	-1.016
	(0.173)***	(0.117)***	(0.142)***
Hunan	-0.589	-0.879	-0.557
	(0.164)***	(0.108)***	(0.156)***
Guangxi	-0.656	-1.364	-1.12
	(0.173)***	(0.130)***	(0.150)***
Guizhou	-1.148	-1.636	-0.904
	(0.176)***	(0.112)***	(0.141)***
Constant	-0.651	0.362	1.321
	(0.163)***	(0.121)***	(0.154)***
Observations	7588	7644	6948
R-squared	0.116	0.205	0.15
Wald tests for equality			
$\beta 1993 = \beta 1997$	F(1, 178) = 34.02***	F(1, 178) = 21.85***	F(1, 178) = 2.34
$\beta 1997 = \beta 2000$	F(1, 178) = 2.50	F(1, 178) = 10.80***	F(1, 178) = 4.29**
$\beta 1993 = \beta 2000$	F(1, 178) = 48.19***	F(1, 178) = 58.94***	F(1, 178) = 12.45***

Robust standard errors, clustered at the community level, in parentheses; no. of clusters is 179
* significant at 10%; ** significant at 5%; *** significant at 1%

Source: China Health and Nutrition Survey, 1991-2000

Note: The sample size is lower for BMI than for weight or height because BMI measures are only valid for children aged two years and older

8. Conclusion

The reduction in the prevalence of malnutrition in China during the 1990s was dramatic and unparalleled. Cross-sectional analysis shows that, from 1991 to 2000, the prevalence of undernutrition among children under 12 fell by about one third while the prevalence of severe undernutrition fell by about one half. Moreover, analysis within cohorts reveals that at least half of the children who were underweight or stunted in 1991 had moved out of this state in subsequent periods of observation.

While not wanting to detract from or diminish these achievements in the least, it would be beneficial to conclude this analysis by underscoring some of the areas of concern that are concealed by these aggregate figures – concerns that only became evident after detailed disaggregation of the patterns and trends. First, there appears to have been a slow-down in the pace of improvement towards the end of the decade; most of the gains occurred in the earlier part of the 1990s and the changes in height-for-age z-scores between 1997 and 2000 were not significant. Second, some of the improvements were regressive, with the greatest gains accruing to those subgroups that were already better-off at the beginning of the 1990s. Third, despite the improvements, in 2000 there remained a sizeable percentage of children still suffering from malnutrition – 15% were underweight and 21% were stunted. Fourth, there is disturbing evidence of a reversal of the long-term trend in three provinces: the prevalence of stunting increased sharply between 1997 and 2000 in Guizhou, Shandong and Henan.

Since the scope of this paper is confined to a documentation of the size, direction and variation of these changes, the obvious next avenue of enquiry is to investigate which factors may have contributed to these improvements. The rapid income growth of the 1990s, and its effect on the purchase of child health inputs, is one obvious candidate. Less well-known, perhaps, are specific policies that could be hypothesized to have affected child health status. The reform of the health care system and inter-temporal variation in the stringency of the one-child family policy may also be salient. It is these questions that will be addressed in the third paper of this dissertation.

Appendix: Additional figures and tables for Paper I

Table 5: CHNS sample size and characteristics, children under 12, 1991-2000

	1991	1993	1997	2000
Total sample	2,457	2,173	1,739	1,275
Urban	642	540	491	314
Rural	1,815	1,633	1,248	961
Male	1,296	1,167	938	697
Female	1,161	1,006	801	578
Under 3 years	447	291	210	179
3-5 years	737	596	263	264
6-8 years	622	677	526	308
9-11 years	651	609	740	524
Jiangsu	243	222	229	173
Shangdong	297	247	167	92
Henan	318	316	302	212
Hubei	400	368	273	195
Hunan	365	294	176	139
Guangxi	441	386	284	208
Guizhou	393	340	308	252

Source: China Health and Nutrition Survey, 1991-2000

Note: Sample sizes shown are for weight-for-age z-scores; the total sample size for height-for-age z-scores is slightly lower (2,435 in 1991, 2,158 in 1993, 1,737 in 1997 and 1,258 in 2000), but exhibits a similar distribution across waves and subgroups.

Table 6: Disparities in the prevalence of underweight and stunting among children under 12, by demographic subgroup, 2000

	Underweight (%)			Stunting (%)		
	Severe	Moderate	Mild	Severe	Moderate	Mild
Total	3	15	45	5	21	50
Urban	1	9	36	4	13	36
Rural	4	17	48	6	24	54
Male	3	13	43	5	21	49
Female	4	17	48	5	21	50

Source: China Health and Nutrition Survey, 1991-2000

Table 7: Disparities in the trends in underweight and stunting among children under 12, by demographic subgroup, 1991-2000

Underweight										
	Moderate					Severe				
	1991	1993	1997	2000	Percentage change 1991-2000	1991	1993	1997	2000	Percentage change 1991-2000
Total	24	20	18	15	-38	6	5	4	3	-50
Urban	18	19	12	9	-50	5	5	4	1	-80
Rural	26	21	20	17	-35	7	5	4	4	-43
Male	23	20	17	13	-43	6	5	4	3	-50
Female	25	20	18	17	-32	7	5	4	4	-43
Stunting										
	Moderate					Severe				
	1991	1993	1997	2000	Percentage change 1991-2000	1991	1993	1997	2000	Percentage change 1991-2000
Total	30	26	20	21	-30	8	7	4	5	-38
Urban	24	20	16	13	-46	8	5	5	4	-50
Rural	32	28	22	24	-25	9	7	3	6	-33
Male	30	26	21	21	-30	9	6	3	5	-44
Female	30	26	19	21	-30	8	7	5	5	-38

Source: China Health and Nutrition Survey, 1991-2000

Table 8: Variation in the prevalence of underweight and stunting, by province, 2000

	Underweight			Stunting		
	Severe	Moderate	Mild	Severe	Moderate	Mild
Jiangsu	0	5	27	1	7	27
Shandong	0	2	10	7	20	34
Henan	3	10	29	8	18	41
Hubei	3	24	58	2	17	55
Hunan	2	11	42	4	18	45
Guangxi	4	13	54	1	18	49
Guizhou	7	28	69	11	40	76

Source: China Health and Nutrition Survey, 1991-2000

Table 9: Pooled probit models of the correlates of stunting and underweight

	Stunting	Underweight
Age (mths)	0.019	-0.001
	(0.007)***	(-0.005)
Age-squared	-0.001	0.001
	(0.001)**	(0.000)***
Male	0.003	-0.015
	(-0.014)	(-0.01)
Urban	-0.103	-0.089
	(0.024)***	(0.018)***
Year dummies		
1993	-0.038	-0.038
	(0.010)***	(0.010)***
1997	-0.093	-0.074
	(0.013)***	(0.012)***
2000	-0.092	-0.101
	(0.014)***	(0.013)***
Province dummies		
Shandong (omitted)		
Jiangsu	-0.019	0.098
	(-0.055)	(0.042)**
Henan	0.007	0.183
	(-0.054)	(0.043)***
Hubei	0.032	0.282
	(-0.057)	(0.043)***
Hunan	0.074	0.2
	(-0.054)	(0.040)***
Guangxi	0.104	0.354
	(0.058)*	(0.045)***
Guizhou	0.266	0.424
	(0.063)***	(0.039)***
Observations	7588	7644
Chi-squared tests of equality		
$\beta1993 = \beta1997$	Chi2(1) = 16.5***	Chi2(1) = 6.74***
$\beta1997 = \beta2000$	Chi2(1) = 0.01	Chi2(1) = 5.19**
$\beta1993 = \beta2000$	Chi2(1) = 13.57***	Chi2(1) = 19.99***

Marginal effects are shown
Robust standard errors, clustered at the community level, in parentheses; no. of clusters is 179
* significant at 10%; ** significant at 5%; *** significant at 1%

Source: China Health and Nutrition Survey, 1991-2000

Note: The sample size is lower for BMI than for weight or height because BMI measures are only valid for children aged two years and older

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PAPER II

ASSETS OR INCOME CONSTRUCTS?

IMPLICATIONS FOR THE ESTIMATION OF CHILD NUTRITIONAL STATUS IN CHINA

1. Introduction

In models of the determinants of child health, an indicator of household income is almost always included as an explanatory variable. Its inclusion is motivated by the theoretical framework of child health production models: for utility-maximizing households, the quantity and quality of market-produced health inputs purchased depends, in part, on the budget constraint[18]. *Ceteris paribus*, an increase in the availability of financial resources increases the investment in health-enhancing goods. Thus, in conditional child health demand models, health is a normal good: children from households with more financial resources are expected to have better nutritional status.

For a variety of reasons, to be explored in this paper, it is disputable whether the income data collected in household surveys in developing countries are a valid and reliable representation of the financial resources that are crucial for child health. Also, many datasets that are rich in health data do not contain quality

[18] See Becker (1965; 1981) for the development of the microeconomic model of household production, Grossman (1972) for the expansion of the model to the demand for health, and Behrman and Deolalikar (1988), Strauss and Thomas (1995) and Currie (2000) for comprehensive reviews of its application to child health in developing and developed countries.

income components[19]. Consequently, in models of child nutritional status, household assets are frequently employed as an alternative to income[20].

Whether income or assets are used can influence conclusions about the effect of economic status on nutritional outcomes in populations. By changing where the researcher ranks the household – or child – in the sample distribution, the choice of indicator alters the observed relationship between economic status and child health outcomes. The extent of this change will depend not only on whether an income measure or an asset measure is used, but also on differences in the way that each is constructed. Moreover, the magnitudes of these differences will vary by cultural environment since the extent of different measurement problems is context-specific.

In this paper, the effect of the choice of indicator on conclusions about the effect of economic status on child nutritional status is explored within the Chinese context. First, potential measurement problems associated with the income measure are discussed. Then, a series of asset indices are constructed and the rankings that each produces are compared to each other and to the rankings produced by income measures. Finally, using income/wealth quintiles constructed from the income measures and asset indices, a series of reduced-form child health demand models are estimated that explore how the use of different asset and income measures alter conclusions about the effects of economic status on child nutritional status. Since the focus is on the appropriate variable to use to capture the budget constraint, the conceptualization of economic status is a narrow one: household economic status is interpreted as a low level of (financial) resources[21].

[19] For example, while the Living Standards and Measurement Surveys (LSMS) contain detailed information on income, consumption and asset ownership (Grosh and Glewwe 1995), the Demographic and Health Surveys (DHS) typically contain data on assets only.

[20] In addition to assets, consumption data are often used to capture household resources in developing countries where they are, in fact, preferred to income data (Ravallion 1998; Deaton and Zaidi 2002).

[21] A broader conceptualization of household well-being may include other dimensions, such as education or autonomy. While these are indeed causal to child health, they are excluded from the conceptualization because they do not enter the model via the budget constraint.

2. Data

Data are drawn from a large-scale household survey conducted in China from 1989 to 2000, the China Health and Nutritional Survey (CHNS). The analysis in this paper will exploit the 2000 wave and cover the seven provinces of Jiangsu, Shandong, Henan, Hubei, Hunan, Guangxi and Guizhou.

2.1 Sampling

In each of the provinces, a multistage, random cluster process was used to draw the sample. Counties in each province were stratified by income (low, middle, and high) and a weighted sampling scheme was used to randomly select 4 counties in each province. In addition, the provincial capital and a lower income city were selected where feasible. Villages and townships within the counties, and urban and suburban neighborhoods within the cities, were randomly selected. In each wave of data collection, newly-formed households residing in sample areas were included and, from 1997, replacement households and replacement communities were added to the sample.

In 2000, the sample consisted of 15,648 individuals in 216 primary sampling units and 1,282 children under twelve for whom anthropometric indicators of nutritional status are available.

2.2 Description of income data

The information on income in the CHNS data is superior to that in most datasets from developing countries. It includes information on wages and salaries[22], income and expenses related to various types of self-employment, including home gardening[23], home farms[24], collective farming[25] activities, animal

[22] In the CHNS, wage-related income refers to the income "of those who are employed by private or collective enterprises and paid on a regular basis, and of those who are engaged in agriculture, animal husbandry, and fishing and are paid on a regular basis" (CHNS 1993: 10). The annual income figures include the monthly salary and any annual bonuses.

[23] Home gardening includes the cultivation of vegetables, fruit and trees either on large plots of land that the household has contracted or on small private plots or yard gardens.

husbandry[26], household fishing[27], and small household business[28]. In addition, information on social security income[29], rental income[30] and remittances is also available. Survey questions elicited information on income in-kind and the market value of home (agricultural and other) production consumed. Data are available at the sub-individual level, i.e. on income and expenses related to *each* type of farming and business enterprise that *every* household member engages in.

2.3 Description of asset data

In the 2000 wave of the CHNS survey, there is asset information on the ownership of radios, VCRs, DVD players, personal computers, black and white TVs, color TVs, telephones, washing machines, refrigerators, air-conditioning, sewing machines, electric fans, microwave ovens, rice cookers, pressure cookers, bicycles, tricycles, motorbikes and cars. Information on certain household utilities or services, namely the source of drinking water, type of fuel used in cooking and type of toilet facility, is also available and is used to construct some of the asset indices in this paper.

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[24] This includes agricultural activities, excluding those included under home gardening.

[25] Here, income from collective fishing, collective animal husbandry and collective agricultural activities includes any income that is not paid in the form of a regular wage or salary. For agriculture, a collective would be a cooperative farm run collectively by several households or individuals who assume responsibility for the profits or losses of the farm and practice collective accounting. For fishing, it could include collective fishing ponds, fishing grounds, or collective fishing in the sea.

[26] Animal husbandry activities include the rearing of pigs, cattle, sheep, horses, mules, rabbits, chickens, ducks, geese etc. as well as silkworms, honey bees etc.

[27] In addition to fishing in ponds, fishing grounds or ocean, fresh water or seawater cultivation of aquatic products is included in this definition.

[28] This refers to small handicrafts and small commercial businesses owned by a household or individual, e.g. carpentry, shoe repair, tailoring and restaurants, but excludes people who work in a small handicraft or small commercial business run by the state, a collective or another individual.

[29] Social security income refers to public transfers such as pensions, unemployment benefits, and other social assistance and social insurance benefits.

[30] Rental income includes cash income obtained from renting out properties, rooms, vehicles, equipment and land.

2.4 Description of data on nutritional status

Nutritional status is operationalized as age- and sex-standardized height-for-age z-scores, constructed on the basis of the 2000 CDC growth charts. Adhering to the fixed exclusion ranges for implausible values recommended by WHO (1995), height-for-age z-scores less than -5 and greater than +3 are excluded from the analysis.

3. Limitations of using income to rank households in China

Income is an intuitive measure of living standards because, in an economy with well-developed markets, having income enables individuals to purchase the goods and services that provide utility. Indeed, it is the preferred measure of living standards in developed countries (Sahn and Stifel 2003). However, the measurement of income is exceptionally difficult, especially in developing countries. In the context of models of child nutritional status, this difficulty relates in part to the issue of construct validity, i.e. how well current income functions as a measure of the concept of resource availability. It is also related to issues of reliability, i.e. how consistently income is measured in the population, both across time and space.

3.1 The permanent income hypothesis/consumption smoothing

Over multiple periods, current income may diverge from consumption and may not accurately reflect actual resource availability. This is because it is possible to save from current income, to finance current consumption by borrowing against future income, and to draw down savings. In addition, households may participate in insurance or risk-pooling arrangements, and households facing a period of low income may receive transfers (remittances) from other households, enabling them to smooth their consumption across high income and low income periods. This is especially likely to be the case for households that face pronounced inter-temporal variability of income, such as seasonal workers, self-employed farmers and business owners who experience a considerable time lag between investment in their farms/businesses and yields, and informal sector workers who face uncertain income streams. One consequence is that income-based measures of living standards may deviate from consumption-based measures. In China, household survey data from the 1990s, for example, showed that income-based measures of living standards provide a more conservative estimate of economic status than consumption measures (World Bank 2001).

The extent to which measured current income will diverge from actual resource availability as a result of consumption-smoothing will depend on a number of factors:

First, the length of the reference period across which income is measured is important. While the permanent income hypothesis is likely to hold over shorter reference periods such as days, months and seasons, households cannot smooth consumption indefinitely. In the long-run, in the absence of inter-generational transfers, income must at least be equal to consumption. Thus, using a cross-sectional survey to capture income over a particular period, for example a month, is unlikely to provide an accurate picture of income during a full year, especially for agricultural households where income is seasonal[31].

Second, the extent to which consumption-smoothing occurs depends on how developed savings, borrowing and risk-pooling institutions – formal or informal – are, and especially the extent to which they operate successfully for the poor whose income may be more variable and whose consumption may be more precarious. In China, the poor do not appear to be too severely credit-constrained. A 1997 survey of six poor counties, for example, showed that nearly two-thirds of households either had an outstanding formal loan or felt they could get one if they wanted it (Park and Ren 2001). The availability of commercial finance is growing and microfinance institutions are mushrooming (Tsai 2004), although there remain many challenges in reaching rural areas[32]. Nationally, China has the highest savings rate in the world: in 2002, savings as a percentage of GDP stood at 42%, half of which was accounted for by Chinese *household* savings (Morrison 2005).

Third, the extent of inter-household transfers of income, which enable households to smooth consumption through drawing on the income streams of other households, is relevant. Inter-household

[31] The thoughtful use of reference periods in the CHNS income questionnaire allows for much of the variability in income to be captured. Wage and salary income, which should be more stable and regular, is measured on a monthly basis. By contrast, agricultural income and small business income, which is more likely to fluctuate due to seasonality effects and local business cycles, is recorded on an annual basis. However, the longer the reference period used, the greater the potential recall bias.

[32] Since much rural land is collectively owned, it cannot be used as collateral; microfinance programs are not permitted to register and take deposits; those who lack credit are increasingly found in remote, mountainous areas (Park and Ren 2001).

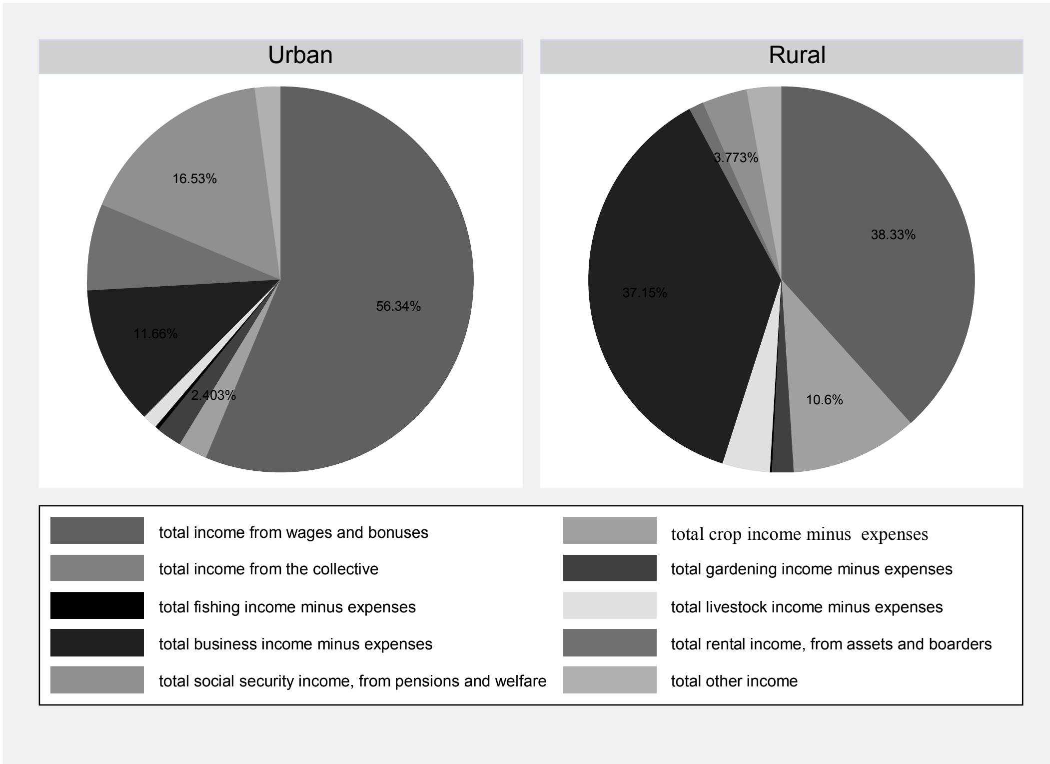
transfers may be very regular and predictable in nature, for example transfers from adult children to elderly parents living outside the household, or they may be more episodic, such as transfers arising from informal obligations to help other households in the community who face sudden need. There is evidence that, for some households in China, the value of inter-household transfers can be fairly high: using data from the 1988 China Household Income Project, Secondi (1997) found that while transfers averaged 34.4 yuan per family per year, they averaged as much as 580.4 among those families that received them.

3.2 Multiple sources of cash and non-cash income

Failure to take into consideration the full range of household income streams will result in a distorted picture of available resources, especially in economies with high levels of self-employment, agricultural work and informal economic activity. In the Chinese case, the distortions would be most pronounced in rural areas. As Figure 10 illustrates, in the seven provinces of China included in this analysis, defining income only as "wages and salaries" would omit 44% of all income flows in urban areas (44%) and as much as 62% of income flows in rural areas. Net[33] income from various types of self-employment, including home gardening, home farms, collective farming activities, animal husbandry, household fishing, and small household business accounts for more than half of all income in rural areas. In urban areas, social security, small business and rental incomes are the next most important, together accounting for more than a quarter of all income.

[33] Each income variable reflects net values. It is constructed by subtracting all expenses associated with a particular income sources from the total revenue from that income source. In order for the pie charts to be constructed, positive mean income values were needed and, so, net income was entered as a zero rather than a negative number when net losses were incurred.

Figure 10: Composition of total household income, by urban-rural area, 2000



Source: China Health and Nutrition Survey, 2000
Note: Figures are based on mean values and reflect income in the previous year

Failure to value income in-kind, i.e. to include flows of goods in addition to flows of money will underestimate the resources that the household has at its disposal to invest in the production of child health. For example, over and above the salary, many Chinese households receive in-kind benefits, such as food and gifts, from work units and local enterprises, or from other households. In addition, and particularly in the earlier years of the survey, yields from collective farming and fishing activities can be an important source of income-in-kind. Finding an appropriate price at which to value these benefits is difficult and an approach frequently adopted in household surveys (including the CHNS) is to ask household heads to estimate the market value of items received. While this approach may reflect the correct monetary equivalent of the items (assuming that the household head can perfectly estimate the items' value), the resultant welfare improvement is unlikely to be equivalent to an income transfer because these items will probably not reflect the utility-maximizing purchasing decisions that households would have made with the equivalent cash income. Consequently, the estimated market value of the items received should ideally be discounted.

There is also a value associated with the consumption of home production. Households engaged in agricultural, animal husbandry or fishing activities may consume some or all of their crops rather than sell them on the market, and may satisfy a large share of their household food needs in this way. The household head's estimate of the price he would have to pay to purchase those items at market is sometimes used to value these items[34], but this strategy is prone to measurement error, especially in communities in which most agricultural production is subsistence or where markets are not well-developed.

3.3 Adjusting for spatial variation in prices

Households that are observed to have the same income may face different prices if they are located in different geographic areas so that the number of items (and thus utility) that can be purchased for a given income varies across households. Consequently, if income is to be used as a measure of resource availability, it needs to be adjusted for prevailing prices. Indeed, in China, Ravallion and Chen (1997) and Démurger *et al*. (2005) have shown that failure to consider cost of living differences leads to substantial overestimation of income inequalities.

3.4 Adjusting for household demographics

Since households can have different sizes, compositions and internal economies of scale, the household level of income may not be an accurate reflection of the resources available to each individual within that household. The simplest approach is to adjust household income for the number of household members, while a slightly more complex approach is to set fixed equivalence scales to account for the differences in income needed by different households to attain the same standard of living[35]. Failure to do so will result in

[34] This is the approach adopted in the CHNS. Household heads were asked to estimate the market value of crops produced and consumed at home, by crop/animal type.

[35] While not the only approach to deriving equivalences scales (see, for example, Deaton and Zaidi 2002), the use of fixed adult equivalence scales is the most common. In this approach, the number of adult equivalents is given by the formula $AE=(A+\alpha K)\theta$ where A is the number of adults in the household, and K is the number of children. The parameter α, lying between 0 and 1, is the cost of a child relative to an adult.

the over-estimation of living standards in larger households, households with a relatively large number of adults compared to children, and households with limited economies of scale in consumption.

3.5 Response bias in income measures

Two types of measurement error that are common to many components of household surveys are particularly pronounced in the income components of surveys: non-response and recall bias.

A reluctance or refusal to disclose income information is to be expected among at least some respondents. This may manifest itself particularly strongly in contexts where income-tests determine access to grants and social services, or where the relatively well-off are expected to fulfill certain financial obligations to the community.

Some (potentially non-random) recall bias is also likely since households are typically asked to recall income received in a previous period, sometimes up to one year prior to the survey. Unless households have detailed household accounting systems, they need to estimate the magnitude of this income and it can be expected that the longer the recall period, the more inaccurate the figures. There is a tendency to round off figures, resulting in the lumping of data around certain “round” numbers. See Appendix Figure 17 for a stark illustration of data lumping in reports of the value of crop sales from the CHNS, for example.

4. Using asset indices to capture changes in economic status

An alternative indicator of resource availability or household economic status is some measure of a household’s asset ownership. Assets reflect a household’s accumulated stock of wealth and, thus, capture a longer-term dimension of the concept of household economic status than the flow concept of income. Ordering households by asset ownership may or may not result in the same ranking as when households are ordered by income. The extent to which this holds depends, first, on how consistently these indicators are measured across different households. It will also depend on the extent to which households are

The parameter θ, also lying between 0 and 1, controls the extent of economies of scale, with lower values reflecting more shared public goods in the household.

homogenous in their marginal propensity to use their income to acquire assets or to save[36]. More generally, different individuals with the same income may have different accumulated wealth stocks if they differ in their motives for wealth accumulation due to life-cycle considerations or precautionary concerns.

With respect to issues of validity and reliability, measures of asset ownership offer some distinct advantages over measures of income. Because assets are directly observable goods and are typically permanent acquisitions, their use avoids many of the problems associated with income measures, such as recall bias, reference periods and seasonality. In poor countries, where assets are fewer, asset information is easier and less expensive to capture than it would be in developed countries. Measurement becomes more complex, however, as one attempts to capture a greater range of assets and the qualitative heterogeneity of assets of a particular type. Also, strictly speaking, a complete asset measure would include the total value of assets *and* liabilities of the household at a point in time, but liability information is seldom available[37].

One way to use assets as a measure of household resources in models of child nutritional status is to construct asset indices[38]. The use of multiple indicators rather than a single asset as a proxy for economic status is likely to increase the validity and the reliability of the results. What distinguishes different types of asset indices from each other is the approach used to assign coefficients, or weights, to the observed variables.

The mathematical representation is of a single aggregate welfare (economic) measure A for each household i, which is created through the explicit or implicit application of weights w_k to each of k asset indicators x_{ik}.

$$A_i = \sum_k \left[w_k x_{ik} \right]$$

[36] In this analysis, assets are interpreted as durable goods, i.e. wealth held in the form of tangible observable goods. Consequently, wealth that is held as financial assets (stocks, bonds, investments) is treated as savings.

[37] The absence of information on liabilities can result in substantial overestimation of living standards, especially in areas with developed housing (and other) finance markets where households may take on large mortgage (and other) debt. The extent to which this affects household rankings depends on the extent to which household ownership of durable goods is correlated with the magnitude of household (and other) liabilities.

[38] The construction of asset indices is not the only way in which information on household assets is used in econometric models of child health. An alternative approach is to include all asset variables separately in the regression model. The estimation procedure would then implicitly create weights on the variables.

Classes of methodologies to assign weights w_k include:

(i) Equal weighting, where $w_k = 1$

(ii) Unequal weighting

(iii) Using the monetary value of the asset as the weight

(iv) Principal components analysis

No single approach can be thoroughly defended as the most appropriate one on theoretical grounds. Nor can these asset indices be used directly for poverty analysis since they have no nominal anchor in a measure of income or expenditure or wealth (Kolenikov and Angeles 2004), with the exception perhaps of the monetary value index (see section 4.3). However, they can be used to rank households and, consequently, they are appropriate for use as a measure of economic status or resource availability in models of child health.

All asset indices discussed in this paper are constructed on a household basis and consist of fourteen household consumer durables, four types of privately-owned transportation and, in the case of some of the indices constructed using principal components methodologies, access to household services[39]. Variables indicating the value of housing and land are excluded because poorly developed housing and land markets mean that a household's housing quality may not closely reflect its economic status. In urban areas of China, this is the legacy of a complex socialist system of housing allocation for employees of the government and of state-owned enterprises and subsidies in the late 1990s reform period (Zhou 2006; Wang *et al.* 2005) and in rural areas of China, as in rural areas of most developing countries, many homes are not purchased but rather constructed using household labor and a mix of purchased and gathered goods so that estimating their monetary value is challenging (Morris *et al.* 2000). More generally, land sales (at market value) are infrequent which, coupled with the fact that land quality is typically highly heterogeneous, means that respondents may struggle to accurately assess land value.

[39] The consumer durables are a radio, television, fan, sewing machine, rice cooker, pressure cooker, refrigerator, washing machine, telephone, vcr or dvd, air conditioner, camera, microwave and computer; the four types of transportation are a car, bicycle, tricycle and motorbike; household services include the source of drinking water, type of fuel used in cooking and type of toilet facility.

4.1 Equal weighting of assets

The simplest asset index is a linear index constructed by the simple aggregation of asset variables, with equal weighting applied to all assets. Each type of asset is included only once, even when households own multiple units of a particular asset. This simple index provides a useful benchmark against which to compare the distributions produced by other asset indices.

$$A_i = \sum_k \left[w_k x_{ik} \right] \text{ where } w_k = 1$$

Since only one unit of each asset type is included in the construction of the simple index, the above methodology may result in household rankings that are counter-intuitive. For example, households with one car and one bicycle may be ranked higher than a household with two cars, all other assets held constant. To account for this, an alternative asset index can be constructed that adjusts for ownership of multiple numbers of the same asset.

$$A_i = \sum_k \left[w_k f_{ik} x_{ik} \right] \text{ where } w_k = 1 \text{ and } f_{ik} \text{ is the number of units of the asset owned by the household}$$

As can be seen in Figure 11, this method yields a wealth distribution that is more right-skewed and has a larger standard deviation than the previous one. By incorporating multiple goods, it distinguishes finer degrees of welfare.

These two methods both suffer from the shortcoming that they fail to take into consideration variations in the value of the different constituent assets and the qualitative heterogeneity among assets of a particular type. In addition, they may be biased by the presence of inferior goods – or more accurately, items that may be inferior for households within particular economic strata. Sewing machines, for example, may be beyond the reach of the poorest households, owned by the middle stratum of non-poor, and not owned by wealthy households who purchase all of their clothing. Still, the fact that individual assets may be weak measures of economic status in particular circumstances does not mean that the index as a whole is a poor reflection of economic status. On the contrary, it is part of the justification of the use of an index rather than a single item to measure economic status.

4.2 Unequal weighting of assets

Assigning equal weights to each asset, as in section 4.1, simplifies the construction of the index, but it is an arbitrary choice which, in addition, fails to consider that assets of a particular type are not homogeneous with respect to their value or quality. An alternative is to apply different weights to different items in the index – a method that is not necessarily any less arbitrary than equal weighting if the determination of the weights is at the discretion of the researcher. To assign weights appropriately requires fairly intimate knowledge of the context. For example, a car would not be as relevant a wealth indicator in big cities with well-developed public transportation infrastructure as it would be in suburban areas. Local preferences are also important, and the type of goods that one acquires as one's economic status improves may be particular to the country, and even to the micro-environment (Krishna 2004), implying that the same goods should be weighted differently in different cultures.

One possibility is to give a greater weight to more luxurious items. For example, a DVD player could be weighted more heavily than a car even though the car is more expensive, if one believes that even middle-class households treat cars as necessities. This approach also helps to overcome the problem of disproportionate importance being given to inferior (substitute) goods since they can be assigned the smallest weight.

As an example of how the application of weights can substantially alter the wealth distribution created by an asset index, slight variation is introduced into the unweighted index developed in the first part of section 4.1 by subjectively assigning weights of 1, 2 or 3 to index components. A higher number indicates an item that is more luxurious, and therefore more likely to be found in wealthy households[40].

[40] The weights chosen are as follows: radio=1, television=1, fan=1, bicycle=1, tricycle=1, sewing machine=1, rice cooker=1, pressure cooker=1, refrigerator=2, washing machine=2, phone=2, motorbike=2, vcr/dvd=3, car=3, air conditioner=3, camera=3, microwave=3 and computer==3. While it may be surprising to the reader to observe a weight of 1 for a television, the fact that more than 90% of households in this sample own at least one television means that it cannot be considered a luxury good.

$$A_i = \sum_k \left[w_k f_{ik} x_{ik} \right] \text{ where } w_k \text{ is an element of } \{1, 2{,}3\}$$

A graphical representation of this weighted index (see Figure 11) shows that, by giving progressively more weight to more luxurious items, the index allows more differentiation between households of different economic strata, especially at the upper end of the distribution. The result is a histogram that is more right-skewed (skewness of 0.77) than the unweighted histogram (0.26), and which more closely resembles the characteristic shape of an income distribution. A greater range of weights is likely to result in even greater differentiation and a flatter distribution.

A more objective alternative is to assign a weight equal to the reciprocal of the proportion of households that own one or more units of each item:

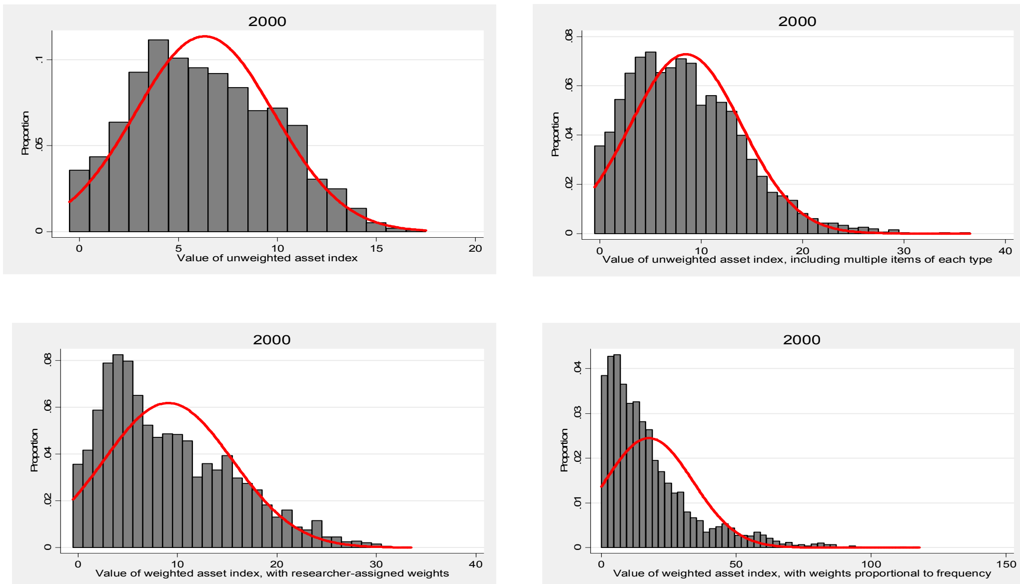
$$A_i = \sum_k \left[w_k \; x_{ik} \right] \text{ where } w_k \text{ is equal to the reciprocal of the proportion of households}$$

that own the item

This weighting system is based on the assumption that households are progressively less likely to own a particular item the higher its monetary value (Morris *et al.* 2000). Therefore, assets that are more likely to be found in homes of higher economic status carry a higher weighting[41]. The presence of inferior goods that are less frequently owned than normal goods would violate this assumption. An example from the CHNS sample is televisions: the more expensive color televisions are more frequently owned than black and white televisions, and if included separately in the index would be assigned a lower weight (1.6) than black and white televisions (2.8). This is one of the reasons why the asset indices constructed in this paper do not to distinguish between the two types of televisions. Like the subjectively weighted asset index, the resultant distribution is more right-skewed than the one that results from the use of the unweighted asset indices.

[41] In this sample, weights range from less than 1.5 for the most commonly owned items, namely televisions, fans and bicycles, to 22 for personal computers and 33 for cars (see Appendix, Table 15).

Figure 11: Wealth distributions resulting from asset indices constructed by equal and unequal weighting methods



Source: China Health and Nutrition Survey, 2000
Note: Asset indices include fourteen household appliances and four types of privately-owned transportation

4.3 Indices reflecting the monetary value of household assets

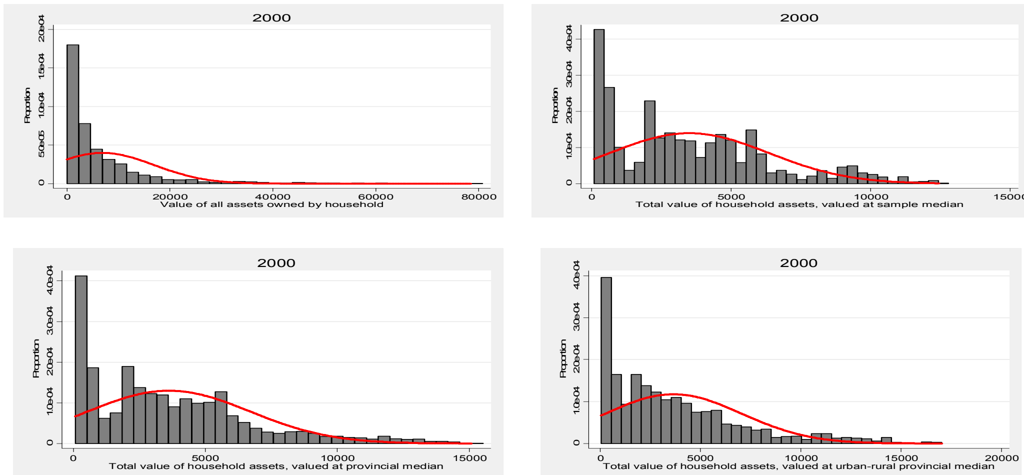
Another approach to constructing an asset index is to estimate the current monetary value of all household assets, using explicit and implicit "prices" as weights in the aggregation. More specifically, the weights can be obtained by either (i) collecting data on the purchase price of each item, date of purchase and suitable depreciation rates, (ii) asking households to estimate the current value of each asset, or (iii) taking the median reported value of the asset across all households as the current market price of an item (Filmer and Pritchett 2001). While data on the purchase price of the household's assets are not available in the CHNS dataset, information on respondents' subjective assessment of the current value of each of their assets allows the consequences of approaches (ii) and (iii) to be explored.

These methods both overcome some of the limitations of the unweighted and weighted asset indices. Most obviously, each asset no longer counts equally. While the subjectively-weighted asset index could have been used to rank assets ordinally in ascending order of value, monetary value methods allow for nominal differentiation. Also, heterogeneity in the quality of assets of the same type is reflected in differential prices. One weakness that these methods share with the weighted methods in section 4.2 is that, because they give proportionately greater weight to items of greater value, depending on the list of assets included, single items of great value may dominate the index.

A shortcoming of approach (ii), which weights each household asset by the respondent's estimate of its current value, is that the weights will be measured with some error due to heterogeneity in respondents' ability to accurately estimate the current value of their assets. For example, the tendency of prices of items with a large technological component to fall over time may result in households *overestimating* the current price of items for which they paid a premium in the past. Households also may not fully understand the effect of depreciation on the current value of their items. On the other hand, they may *underestimate* the value of their items in an inflationary environment where escalating prices of new goods also chase up the value of second-hand goods. Approach (iii), which uses the *sample median* reported value of each asset's value as the weight (i.e. current price) in the index, may alleviate some of the measurement error that arises from household's limited ability to accurately assess the value of their items and may also lessen the impact of single valuable items on the index. As can be seen in Figure 12, using the sample median prices compresses the distribution and reduced the skewing.

Since there may be significant spatial variation in the prices faced by households, it is also informative to examine how the distribution changes if, instead of calculating a sample median, the median value of each asset for each province is calculated. A further step is to calculate the median value of each asset separately for urban and rural areas within each province. As can be seen from the graphs of these distributions (Figure 12), these approaches introduce some variation into the median price (i.e. weight). As the geographic unit of analysis becomes smaller (i.e. as one moves from analysis at the sample level to the provincial level to urban-rural strata within provinces), the range of monetary values increases to reflect the greater variation in median prices.

Figure 12: Wealth distributions resulting from asset indices constructed by weighting by the reported monetary value of assets



Source: China Health and Nutrition Survey, 2000
Note: Asset indices include fourteen household appliances and four types of privately-owned transport; see Appendix, Table 16, for actual median prices

4.4 Constructing an asset index using principal components analysis

Although principal components analysis has been used since the early 1900s (first, by Pearson 1901, and later in a more developed form by Hotelling 1933), and has had diverse applications in the physical, natural, medical, behavioral and social sciences, its use as an alternative means of deriving the weights needed to construct a linear asset index only gained popularity following Filmer and Pritchett's (1998; 2001) papers on the relationship between household wealth and school enrollment in India. Their approach has subsequently been used to measure economic status in studies of the determinants of different human development outcomes, such as education (Filmer 2000; Filmer and Pritchett 1999) and health (Pande and Yazbeck 2002).

Principal components analysis is a statistical technique that transforms an original set of variables into a smaller set of variables (called principal components) that are a linear combination of the original variables, are uncorrelated (orthogonal), and capture most of the variation in the original variables in the smallest number of dimensions (Jolliffe 2002).

The motivation behind a principal components approach to the construction of an asset index is the same as other methods of creating asset indices: to create "a proxy for something unobserved: a household's long-run economic status" (Filmer and Pritchett 2001). Individually, each asset is a weak and mis-measured indicator of household economic status, and, together, they are highly inter-related. Since principal components analysis exploits the variation and reduces the redundancy in these assets, the asset index derived using this approach could be a more efficient proxy for household economic status than indices derived by other methods.

Formally, then, suppose there is an unobserved variable ξ (such as household economic status or resource availability) with a set of k observed indicators x_{1i} to x_{ki} (such as household assets) for each household i. Then, the principal components approach specifies each asset/indicator normalized by its sample mean and standard deviation, and expresses each variable/indicator/asset as a linear combination of a set of underlying components for each household i. Then it finds the linear combination of the variables with the maximum variance – called the first principal component – followed by the second linear combination of variables, orthogonal to the first, with maximum remaining variance, and so forth through subsequent components. While second and subsequent principal components are extracted, in empirical work the first principal component is typically considered to be a satisfactory measure of welfare. The first principal component (expressed in terms of the original unnormalized variables) is, therefore, an asset index (A_i) for household i:

$$A_i = \sum_k \left[f_k \frac{(x_{ik} - \bar{x}_k)}{s_k} \right]$$

where x_{ik} is the value of asset k for household i, $\bar{x}$ is the sample mean, and s_k is the sample standard deviation. The asset index is, thus, equivalent to the sum of the included variables, weighted by the elements of the first eigenvector (i.e. the first principal component).

Principal components analysis is a linear procedure and its theoretical results, including the consistency of the estimates of the factor loadings, are derived under the normality assumption (Kolenikov and Angeles 2004). Consequently, the technique should be used with data that are continuous and approximately normal. Most asset data, however, are discrete – typically binary variables (in the case of ownership of household durables) or ordinal (type/quality of toilet facility). Depending on the nature of the data, and the construction of asset variables, one of three different approaches to PCA may be more defensible.

(i) PCA with dummy variables: The most common approach in PCA studies of asset indices is to enter all asset variables as dummy variables. Consequently, asset variables that are in ordinal form in the original data are recoded. For example, a "source of drinking water" variable with four categories is recoded into four separate dummy variables. However, as pointed out by Kolenikov and Angeles (2004), this approach introduces a lot of spurious correlation: the dummy variables produced from the same ordinal variable are negatively correlated (although the strength of the dependence declines as the number of categories increases) so that the PCA procedure cannot determine whether the main source of the common variation is due to the correlation with the unobserved welfare measure (i.e. household economic status or resource availability), as desired, or due to correlation among the (dummy) variables derived from the same ordinal variable[42]. The result is a noisier covariance matrix, the underestimation of the share of the variance explained by the first few components, and deteriorating goodness of fit measures. In addition, generating dummy variables discards any of the ordinal information that may have been available in any original ordinal variables. The dummy variable approach does have certain advantages, though. One is that it makes no assumptions about the true or natural underlying ordering of ordinal variables. For example, it does not assume that using electricity for cooking is necessarily "better" than using kerosene.

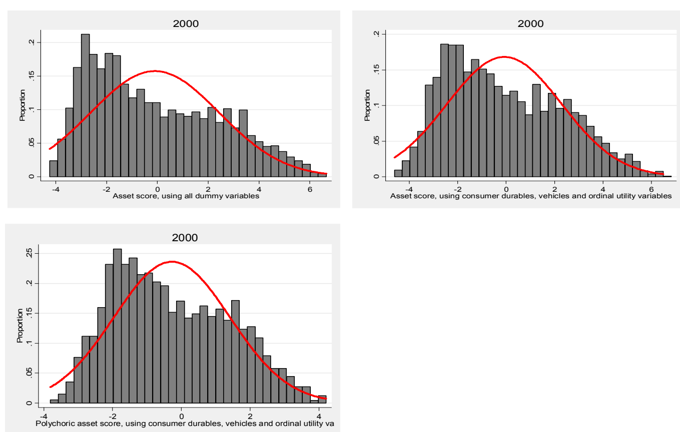
(ii) *PCA with ordinal variables*: This approach is similar to the dummy variable approach except that, where possible, an ordinal construction of the variables is used[43]. As with the dummy variable approach, principal components analysis is performed as if the variables are continuous. Although an improvement on the dummy variable approach for reasons discussed above, this approach still violates the distributional assumptions of principal components analysis and the data are likely to exhibit a high degree of skewness and kurtosis. Correlations will be smaller than with continuous data.

[42] A related point is that in recoding ordinal variables to dummy variables, an arbitrary zero weight is assigned to the omitted category when, in fact, the omitted category could perhaps signify membership of the middle of the wealth distribution.

[43] In other words, variables capturing household utilities and services are constructed so that a higher number indicates a higher level of economic status – subjectively determined by the researcher. For example, the source of drinking water can be coded as 3 for in-house tap water, 2 for in-yard tap water and 1 for well water.

(ii) *PCA using polychoric correlations*: A third approach is to perform principal components analysis on the polychoric correlation[44] ρ between ordinal (or even dichotomous) asset variables. The use of a polychoric correlation assumes that each asset variable reflects an underlying continuous variable with a bivariate normal distribution, and the polychoric correlation extrapolates what the categorical variables' distributions would have been if continuous, adding tails to the distribution. In a recent paper, Kolenikov and Angeles (2004), using a large Monte Carlo simulation, found that the only convincing reason to prefer polychoric PCA over ordinal PCA, though, was the proportion of the explained variance, which was reported consistently only by the polychoric method[45] (see next section). The household rankings – of central interest in this paper – produced by the two methods were very similar.

Figure 13: Wealth distributions resulting from asset indices constructed by dummy variable, ordinal and polychoric PCA approaches



Source: China Health and Nutrition Survey, 2000
Note: All three asset indices include fourteen household appliances and four types of privately-owned transport. In the case of the ordinal (top right) and polychoric approaches (bottom left), ordinally-coded utility variables are also included. For comparison with the distributions of other asset indices, ordinal and polychoric PCA distributions using only the consumer durables and privately-owned transport are shown in the Appendix, Figure 18.

[44] The polychoric correlation is the maximum likelihood estimate of the underlying bivariate normal distribution of two ordinal variables, each created by discretizing an unobserved normally distributed underlying variable.

[45] The ordinal approach and, especially, the dummy variable approach produce estimates of the proportion of the explained variance that are lower than those obtained using the polychoric approach.

5. The performance of different asset indices and income measures

Since the CHNS survey lacks a consumption module, there is no *external* benchmark against which to compare the performance of asset indices and income measures. Nevertheless, it is still possible to gauge the internal performance of the asset measures and to compare the rankings produced by different asset indices to each other and to those produced by income measures.

5.1 Performance of asset measures

5.1.1 Internal goodness of fit of PCA asset measures

An internal goodness of fit measure for asset indices constructed by the PCA approaches is the proportion of the variance explained by the first component (eigenvalue). In all cases, the size of the first component is very large relative to subsequent components, suggesting that the asset score is a very good approximation of the total variation in the sample, with other components consisting largely of noise (see scree plots in the Appendix, Figure 19).

The results presented in Table 10 show that the performance of PCA improves when (i) variables are coded as ordinal rather than as discrete variables, and (ii) when polychoric rather than ordinal PCA, or ordinal rather than dummy variable PCA is used. It also improves when (iii) the number of variables included increases, provided that these variables are ordinal and not discrete. In the PCA with dummy variables, for example, the proportion of variance explained falls from 29.7% to 17.5% as the number of components increases from 14 to 36[46]. By contrast, in the PCA with ordinal utility variables, the goodness of fit measures actually improves when the utilities are added because the methodology takes advantage of the extra information contained in the ordinal coding of those variables. More importantly, the proportion of variance in the sample explained by the first principal component is higher in the ordinal approach than

[46] There are 14 components when the asset index consists of only consumer durables, 18 when the index consists of consumer durables and vehicles and 36 when it consists of consumer durables, vehicles and utilities.

in the dummy variable approach, and much higher still in the polychoric approach[47]. Focusing on the results of the full model where all asset variables are included (i.e. the last row in Table 10), the proportion of the variance explained by the first principal component increases from 17.5% (in the PCA model with all dummy variables), to 25.2% (in the PCA model with ordinal utilities) to 41.5% (in the polychoric model with ordinal variables).

Table 10: Internal goodness of fit: the proportion of variance explained by the first component in 2000

Variables in model	PCA with all dummies	PCA with ordinal utilities	Polychoric PCA	
	Binary durables; binary utilities	*Binary durables; ordinal utilities*	*Binary durables; binary utilities*	*Binary durables; ordinal utilities*
Only consumer durables	29.7%	29.7%	51.8%	51.8%
Consumer durables and vehicles	23.9%	23.9%	42.1%	42.1%
Consumer durables, vehicles and utilities	17.5%	25.2%	Error	41.5%

Source: China Health and Nutrition Survey, 2000

Note: All consumer durable and vehicle variables are dummy variables in the original data; the "error" means the likelihood is ill-behaved during the maximization procedure and, so, cannot be calculated.

5.1.2 Internal coherence of asset indices

Following Filmer and Pritchett (2001), the internal coherence of a particular asset index or income measure can be analyzed by comparing how average ownership of select assets varies across its tertiles. In all the approaches taken to the construction of asset indices, the percentage of households owning these (normal) goods increases as one moves from the lowest to the highest asset tertiles, suggesting that all measures are internally coherent (see Appendix, Table 17 and 18). This method is particularly illuminating in the case of PCA where it shows that the distribution of asset ownership across quintiles is almost identical under the ordinal and polychoric approaches, highlighting the similarity of the rankings produced by these indices.

[47] Similar results have been obtained using datasets from Russia and Bangladesh (Kolenikov and Angeles 2004). In addition, these authors found that the extent to which the dummy variable approach underestimates the explained variance increases as the number of categories in the variables increases.

5.1.3 Correlation among different asset indices

Tests of association highlight differences in the rankings that are produced by different asset indices (see Table 11). The strongest association is found between the indices constructed using the ordinal and polychoric PCA approaches (Spearman value of very close to 1.00), providing further evidence that the computational costs of the polychoric approach may outweigh the (performance) returns – at least for the purpose of ranking households. Both are also very strongly associated with the index constructed using dummy variable PCA (Spearman correlation of 0.97). The three measures that use different median prices (sample, provincial, urban-rural provincial) are also very closely correlated (Spearman correlation coefficients in excess of 0.97). From a practical standpoint, this calls into question the additional value of calculating separate provincial or provincial-urban-rural asset indices, or even collecting detailed price data at this level. Another finding is that the index based on households' reports of the monetary value of their assets, which theoretically, should be a very close approximation of the household's actual economic status, is the most poorly correlated with all other measures[48]. This may mean that the other measures are poor proxies of economic status (or resource availability) or it may mean that households are not very good judges of their own welfare[49].

[48] It is even poorly correlated with other indices that attempt to capture the value of assets, such as the three measures that use median prices (0.63). The closest approximation to it is the index where weights are assigned proportional to frequency of the ownership of each item in the sample (0.75).

[49] It is also possible that the relatively large number of missing values on household reports of the monetary value of goods may be influencing the results. Since these values are not missing at random – for example, the mean per *capita* income in households with missing values is 1,151 yuan compared to 1,204 yuan in households without missing values – this may partially explain the poor correlation.

Table 11: Spearman correlations between different non-PCA asset indices asset measures

Weighting used	1	2	3	4	5	6	7	8	9	10	11
1 Unweighted	1.00										
2 Unweighted, multiple assets	0.93	1.00									
3 Researcher-assigned weights	0.98	0.91	1.00								
4 Weights proportional to ownership	0.93	0.87	0.95	1.00							
5 Monetary value	0.84	0.81	0.87	0.86	1.00						
6 Sample median prices	0.88	0.79	0.91	0.82	0.83	1.00					
7 Provincial median prices	0.88	0.81	0.91	0.82	0.84	0.98	1.00				
8 Provincial urban-rural prices	0.88	0.81	0.90	0.81	0.84	0.97	0.98	1.00			
9 Dummy variable PCA	0.91	0.80	0.90	0.82	0.76	0.87	0.86	0.87	1.00		
10 Ordinal PCA	0.95	0.86	0.95	0.87	0.81	0.90	0.89	0.90	0.97	1.00	
11 Polychoric PCA	0.96	0.87	0.95	0.87	0.81	0.90	0.89	0.90	0.97	1.00	1.00

Source: China Health and Nutrition Survey 2000
Note: All correlations are statistically significant at the 1% level.

5.1.4 Tertile "misclassification" rates of asset indices

Estimating the percentage of households that are classified into different asset tertiles when different methodologies of constructing asset indices are used is one way to test the robustness of the results. High "misclassification"[50] rates would indicate that the choice of asset index will make a large difference to estimates of the effect of economic well-being on child nutritional status. For convenience, the simple unweighted asset index is used as the comparison case.

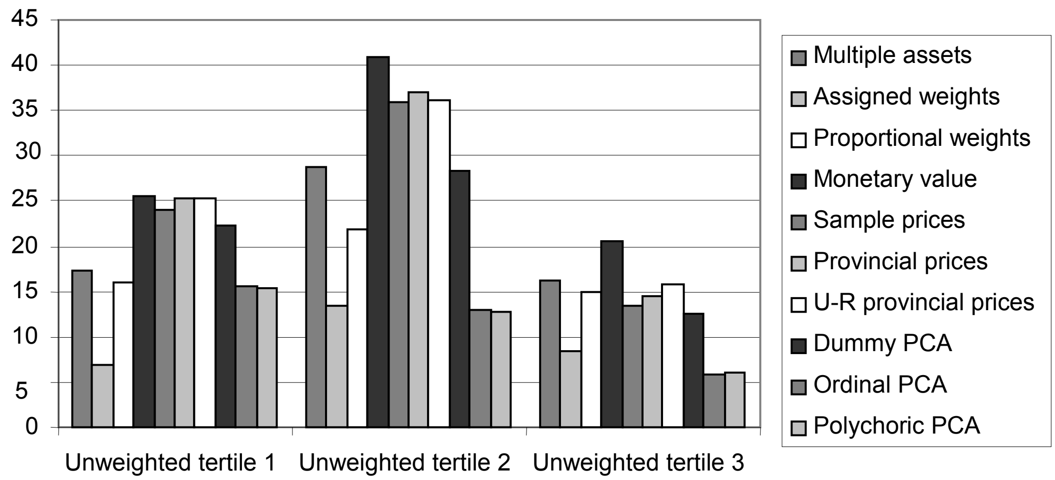
As expected, the highest rates of misclassification occur in the middle tertile (see Figure 14). Households in the top (wealthiest) tertile 3 are more consistently classified across indices than households in the lowest (poorest) tertile 1. Although the incidence of misclassification is fairly high, the magnitude of misclassification is not too large, at least for better-off households: of those households in the top tertile in the simple unweighted index, not a single household was classified in the lowest tertile by any of the alternative asset indices. The classifications that most closely resemble the simple index are those resulting from the ordinal and polychoric PCA approaches, despite the fact that they include an additional set of variables (namely household utilities). The classifications are especially similar in the upper tertile.

Clearly, the choice of index does matter. The fact that the extent of misclassification is greatest for the lowest economic status group (tertile) suggests further that the measure matters most for those who are least well-off and whose children are most likely to be in poor health. The choice of measure not only

[50] The term "*mis*classification" is something of a misnomer since assets are not being compared to some external benchmark, such as consumption, but rather to each other.

affects household ranking, but the effect on ranking is likely to be correlated with the incidence of malnutrition.

Figure 14: Tertile misclassification rates: percentage of households in the unweighted tertiles that are classified in different tertiles using alternative asset indices.



Source: China Health and Nutrition Survey, 2000
Note: All tertiles are compared to the simple unweighted asset index. For example, the first bar shows that 17.28% of the households that were classified as in tertile 1 using the unweighted simple asset index were classified in a different tertile by the unweighted index that includes multiple units of each asset.

5.2 Performance of income measures

Three income measures, chosen because they highlight some of the measurement issues related to income discussed in section 3, are considered:

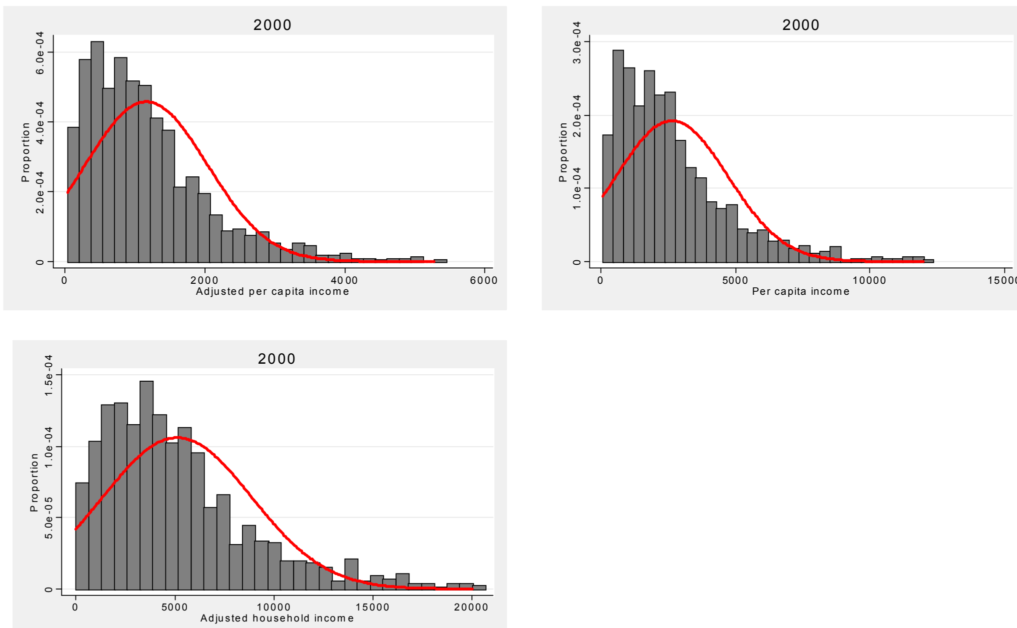
(i) an "adjusted *per capita* income" measure that includes all sources of income discussed in section 3.2, including cash income, income in-kind and the value of home production, and is adjusted for the number of household members and community-level variation in prices;

(ii) a "*per capita* income" measure, that is constructed in the same manner as (i), but is not adjusted for spatial variation in prices;

(iii) an "adjusted household income" measure, that is constructed in the same manner as (i), but is not adjusted for household size;

All measures are constructed to reflect annual values.

Figure 15: Distributions of adjusted *per capita* income, *per capita* income and adjusted household income, 2000



Source: China Health and Nutrition Survey, 2000

Evidence of the internal coherency of all measures is found in comparisons of the mean percentage of households that own a particular asset across the income tertiles of each income measure: as one moves from a lower tertile to a higher tertile, the mean percentage of households owning a particular asset increases. The measures are also highly correlated, with Spearman coefficients higher than 0.9. There is almost no difference in the ranking produced by the *per capita* income measure that is adjusted for spatial variation in prices and the *per capita* income measure that is not adjusted for price variations (Spearman's ρ = 0.996). The correlations between these two measures and the adjusted household measure are equivalent to about 0.91, though, showing that failure to account for household size alters rankings somewhat. An examination of tertile "misclassification" rates does not yield any remarkable differences across income measures.

On balance, then, it appears that using any one of these three measures should not make too much difference to the ranking of households or the estimated effects of economic status on child health.

5.3 Comparison of asset indices and income measures

Even the simplest tests show that the household rankings produced by asset indices and income measures differ substantially. Regardless of the asset or income measure used, the correlation between income and asset constructs is fairly low. In not one case does Spearman's ρ exceed 0.5.

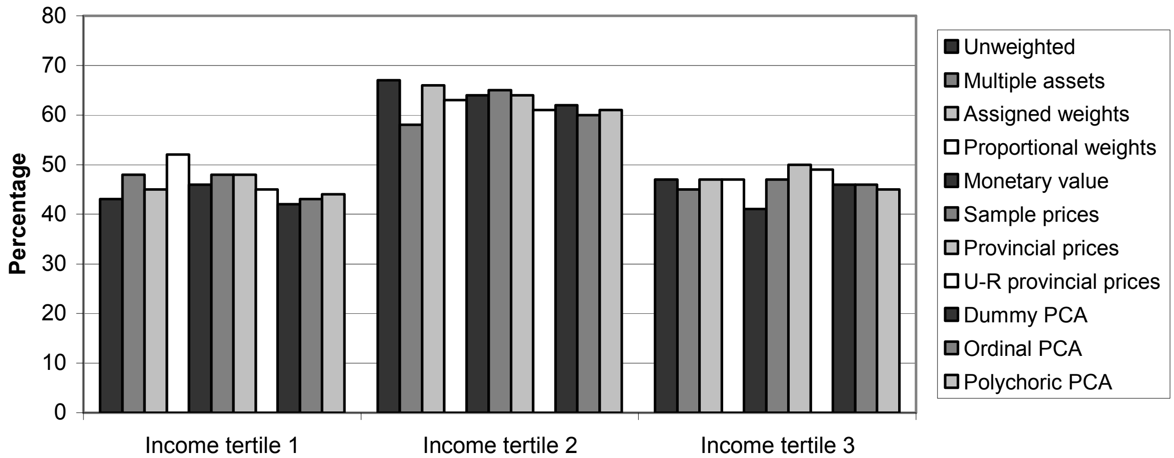
Table 12: Table of Spearman correlations between different non-PCA asset indices asset measures

Weighting used	*Per capita* income (adjusted)	*Per capita* income (unadjusted)	Household income (adjusted)
1 Unweighted	0.442***	0.466***	0.415***
2 Unweighted, with multiple assets	0.415***	0.437***	0.436***
3 Researcher-assigned weights	0.43***	0.451***	0.398***
4 Weights proportional to ownership	0.404***	0.421***	0.384***
5 Self-reported monetary value	0.443***	0.46***	0.417***
6 Sample median prices	0.414***	0.434***	0.365***
7 Provincial median prices	0.414***	0.434***	0.362***
8 Provincial urban-rural prices	0.447***	0.470***	0.391***
9 Dummy variable PCA	0.47***	0.5***	0.411***
10 Ordinal PCA	0.465***	0.494***	0.418***
11 Polychoric PCA	0.465**	0.494***	0.421***

Source: China Health and Nutrition Survey, 2000
Note: All correlations are statistically significant at the 1% level.

A table of tertile misclassification rates which compares the household ranking that results from the use of an adjusted *per capita* income measure to the rankings that result from the use of asset measures shows that that there is a very high rate of misclassification across all income tertiles (Figure 16). The highest misclassification occurs, as expected, in the middle tertile where the percentage of alternately classified children exceeds 60% for most asset indices. Households classified in the lowest tertile by the income measure are not any more likely to be classified in other tertiles than households in the upper tertile are. Moreover, although it is not shown in this graph, a fairly high percentage of households that are classified in tertile 1 by the income measure are classified in tertile 3 by the asset measures and *vice versa*.

Figure 16: Tertile "misclassification" rates: percentage of households in the each income tertile that are classified in different tertiles using the different asset indices.



Source: China Health and Nutrition Survey, 2000
Interpretation: For example, the first bar shows that 43% of the households that are classified in the first tertile using the adjusted *per capita* income measure are classified in a different tertile by the unweighted asset index.

6. The effect of using different constructs on conclusions about the relationship between economic status and child nutritional status

The results presented so far have shown that any particular asset measure ranks at least a proportion of households differently from income measures and from other asset measures. If the differences are large, it is likely that the choice of construct will affect the estimated relationship between economic status and child nutritional status and conclusions about the health disparities between rich and poor.

6.1 The model

To explore this, a conditional child health demand function is estimated. Empirically, current nutritional status (H) is modelled as a function of a vector (Z_1) of current period child-specific characteristics (namely age and sex) and a vector (Z_2) of household-specific characteristics (namely mother's education, mother's height and the financial resources of the household).

$$H = \beta_0 + \beta_1 Z_1 + \beta_2 Z_2 + u \qquad (1)$$

The dependent variable, nutritional status, is captured by the age- and sex-standardized height-for-age z-score[51]. The mean z-score is 1.98, which means that the average child in the sample is mildly malnourished, and just over one-fifth of the sample is stunted[52].

Household financial resources are captured by the income measures and asset indices discussed in the preceding sections. They enter the model as a set of quintile dummy variables which reflect the child's relative position in the sample distribution of each particular index; the poorest children are in quintile 1 and the wealthiest children are in quintile 5 (the omitted category).

Other covariates include the sex of the child which, since data are already age- and sex-standardized by the use of z-scores, will capture social rather than biological influences on differences in the growth patterns of boys and girls. Height-for-age z-scores are modelled as a quadratic function of age (in years) to allows for the cumulative effect of growth-retardation over time. Mother's height, measured in centimetres, is included to capture the child's unobserved growth potential or genetic endowment. Mother's education, which can affect child nutritional status via its effect on income, access to information and enhancement of literacy and numeracy skills (see, for example, Thomas *et al.* 1991; Glewwe 1999; Handa 1999; Christiaensen and Alderman 2004) is represented by a set of dummy variables for 1 to 3 years of primary education, 4 to 6 years of primary education, 1 to 3 years of middle education, and 1 to 3 years of upper middle education or more. The omitted category is mothers who have received no schooling.

6.2 The results

Three models with income measures and eleven models with asset measures are estimated. Because there is no consumption benchmark against which to compare the performance of the asset measures or income

[51] Recumbent length-for-age, rather than height-for-age, is used for children under the age of 36 months because children under the age of three were measured using a length-board.

[52] The term "severely stunted" is used when the height-for-age measurement is less than three standard deviations below the reference median, "stunted" is used when the measurement is less than two standard deviations below the reference median, and "mildly stunted" is used when the measurement is less than one standard deviation below the reference median.

measures, there is no theoretical reason to prefer one set of estimated results to the other. Nevertheless, the different models do illustrate the extent to which the choice of measure can influence the estimated effect of economic status on child nutritional status.

6.2.1 Income measures

The most obvious finding is that all income models lead to the conclusion that increased financial resources are positively associated with improvements in nutritional status (see Table 13). Children in the lowest quintile have height-for-age z-scores that are significantly lower than those of children in the highest quintile, regardless of whether income is adjusted for price variation and for household size.

The magnitudes of the effect of income and, thus, the effect of income disparities on nutritional status, are very similar in the *per capita* income model that is adjusted for price variation (model I) and the un-adjusted *per capita* measure (model II). By contrast, the model that adjusts for price variation but not household size (model III) fails to yield significant quintile coefficients that increase progressively in the magnitude of their effect with each successive quintile. It does show that being poor (in quintile 1) is significantly associated with lower height-for-age z-scores, but the coefficient is 0.1 to 0.15 of a standard deviation smaller than in the other income models.

Table 13: OLS regressions of height-for-age z-scores on income quintiles

	Model I	Model II	Model III
	Per capita income (adjusted)	*Per capita* income (unadjusted)	Household income (adjusted)
Sex (Male=1)	-0.09	-0.092	-0.083
	-0.075	-0.074	-0.075
Age (years)	0.04	0.039	0.046
	-0.074	-0.074	-0.074
Age squared	-0.001	-0.001	-0.001
	-0.005	-0.005	-0.005
Maternal education (“no formal education” is omitted)			
1-3 yrs primary	-0.256	-0.248	-0.499
	-0.247	-0.25	(0.122)***
4-6 yrs primary	-0.251	-0.252	-0.294
	-0.224	(0.123)***	-0.222
1-3 yrs lower middle	0.131	0.127	-0.301
	-0.22	-0.222	-0.216
1-3 yrs upper or more	0.345	0.33	0.398
	-0.223	-0.225	(0.221)*
Mother’s height (cm)	0.049	0.049	0.05
	(0.011)***	(0.011)***	(0.011)***
Quintiles (5th quintile – wealthiest – is omitted)			
1 (poorest)	-0.626	-0.668	-0.529
	(0.130)***	(0.136)***	(0.133)***
2	-0.514	-0.464	-0.244
	(0.123)***	(0.122)***	-0.25
3	-0.457	-0.5	-0.247
	(0.124)***	-0.224	(0.120)**
4	-0.368	-0.335	0.162
	(0.121)***	(0.118)***	(0.128)**
Constant	-8.417	-8.391	-8.642
	(1.700)***	(1.706)***	(1.686)***
Observations	944	944	944
R-squared	0.179	0.181	0.173

Source: China Health and Nutrition Survey, 2000

Note: * significant at 10%; ** significant at 5%; *** significant at 1%; robust standard errors, clustered at the community level, in parentheses; “adjusted” refers to adjustment for spatial variation in prices.

6.2.2 Asset measures

All models deliver a consistent message about child health disparities between the rich and the poorest of the poor: in each case, the mean z-scores of children in the lowest (poorest) quintile are significantly lower than children in the upper quintile (see Table 14). Moreover, the range of the estimated effect for quintile 1 is very narrow: the coefficients on nine out of the eleven models are within one-tenth of a standard deviation of each other, ranging from -0.72 to -0.82. The outlier at the upper end is the asset index that allows for ownership of multiple items of each asset type (model II). Here, the children in the lowest quintile have height-for-age z-scores that are, on average, an entire standard deviation (-1.0) lower than that of children in the upper quintile. The outlier at the lower end at the opposite end is the model in which children are allocated into quintiles on the basis of the household head's assessment of the monetary value of assets (model V). Here, the estimated effect on nutritional status of being in the lowest quintile rather than the highest quintile is only half a standard deviation (-0.50), but when this figure is adjusted for the relative prices of the index components and spatial variation in those prices (see models VI through VIII), the coefficient is brought within the -0.72 to -0.82 range. The three PCA approaches (models IX through XI) resulted in only slightly different coefficients, within 0.05 of a standard deviation of each other.

Looking beyond differences in nutritional status between the rich and the very poor, however, one does not observe the linear relationship between increases in financial resources and improvements in nutritional status that is seen in the *per capita* income measure used in the previous section. Indeed, there is not a single asset or income measure where the mean z-score increases progressively and significantly across the quintiles.

Table 14: OLS regressions of height-for-age z-scores on quintiles constructed from asset indices

	Assets I	Assets II	Assets III	Assets IV	Assets V	Assets VI	Assets VII	Assets VIII	Assets IX	Assets X	Assets XI
	Simple	No. of assets	Assigned	Prop weights	Monetary Value	Sample prices	Province prices	U-R province prices	Dummy PCA	Ordinal PCA	Poly-choric PCA
Sex (Male=1)	-0.108	-0.117	-0.117	-0.124	-0.115	-0.071	-0.073	-0.063	-0.103	-0.102	-0.103
	-0.073	-0.072	-0.074	(0.074)*	-0.079	-0.075	-0.076	-0.074	-0.074	-0.074	-0.074
Age (years)	0.057	0.06	0.054	0.05	0.044	0.075	0.079	0.074	0.053	0.058	0.056
	-0.074	-0.074	-0.075	-0.076	-0.076	-0.077	-0.077	-0.075	-0.072	-0.073	-0.073
Age squared	-0.003	-0.003	-0.002	-0.002	-0.002	-0.004	-0.004	-0.004	-0.002	-0.003	-0.003
	-0.005	-0.005	-0.005	-0.005	-0.005	-0.005	-0.005	-0.005	-0.005	-0.005	-0.005
Maternal education (“no formal education” omitted)											
1-3 yrs primary	-0.311	-0.34	-0.504	-0.413	-0.497	-0.654	-0.296	-0.293	-0.396	-0.342	-0.711
	-0.26	-0.25	(0.138)***	(0.144)***	(0.154)***	-0.261	(0.142)***	-0.253	-0.249	-0.254	(0.137)***
4-6 yrs primary	-0.286	-0.449	-0.29	-0.26	-0.22	-0.244	-0.242	-0.282	-0.38	-0.368	-0.364
	-0.219	-0.201	-0.225	(0.162)*	-0.23	-0.229	-0.222	(0.140)***	-0.223	(0.142)**	-0.221
1-3 yrs lower middle	-0.018	-0.44	0.007	-0.106	-0.076	0.026	0.028	-0.341	-0.082	-0.083	-0.313
	-0.215	(0.140)***	-0.22	-0.131	-0.228	-0.229	(0.131)**	-0.21	(0.131)*	-0.21	-0.213
1-3 yrs upper or more	0.16	0.173	0.206	0.276	0.318	0.167	0.188	0.076	0.053	0.067	0.062
	-0.222	-0.199	-0.224	-0.243	-0.233	-0.232	-0.225	-0.211	-0.222	-0.218	-0.221
Mother’s height (cm)	0.043	0.04	0.044	0.044	0.038	0.041	0.039	0.04	0.048	0.046	0.046
	(0.010)***	(0.010)***	(0.011)***	(0.011)***	(0.011)***	(0.010)***	(0.010)***	(0.010)***	(0.011)***	(0.011)***	(0.011)***
Quintiles (5th quintile omitted)											
1 (poorest)	-0.817	-1.007	-0.735	-0.633	-0.496	-0.727	-0.78	-0.813	-0.764	-0.797	-0.8
	(0.134)***	(0.149)***	(0.150)***	(0.165)***	(0.156)***	(0.138)***	(0.144)***	(0.144)***	(0.148)***	(0.146)***	(0.151)***
2	-0.589	-0.655	-0.324	-0.319	-0.286	-0.254	-0.643	-0.886	-0.69	-0.725	-0.347

	(0.126)***	(0.146)***	-0.264	-0.279	-0.271	(0.134)***	-0.263	(0.138)***	(0.132)***	(0.141)***	-0.258
3	-0.256	-0.294	-0.375	-0.271	-0.385	-0.594	-0.514	-0.523	-0.335	-0.339	-0.344
	(0.142)*	(0.140)***	(0.165)**	-0.239	(0.137)***	(0.135)***	(0.135)***	-0.213	(0.147)**	-0.219	(0.147)**
4	-0.298	-0.037	-0.211	0.046	0.098	-0.218	-0.33	-0.051	-0.238	-0.341	-0.088
	(0.124)**	-0.194	(0.119)*	-0.232	-0.15	-0.135	-0.222	(0.130)***	-0.212	(0.125)***	(0.124)**
Constant	-7.34	-6.83	-7.518	-7.698	-6.745	-7.114	-6.863	-6.88	-8.052	-7.687	-7.728
	(1.668)***	(1.633)***	(1.713)***	(1.748)***	(1.665)***	(1.646)***	(1.663)***	(1.578)***	(1.709)***	(1.716)***	(1.725)***
Observations	910	910	910	910	774	887	887	887	901	900	900
R-squared	0.195	0.213	0.183	0.179	0.162	0.18	0.178	0.193	0.198	0.2	0.2

Source: China Health and Nutrition Survey, 2000

Note: * significant at 10%; ** significant at 5%; *** significant at 1%; Robust standard errors, clustered at the community level, in parentheses

7. Conclusion

The fact that the inclusion of a measure of household financial resources in models of child nutritional status is near-universal, while tests of the robustness of results to different measures of economic status or resource availability are relatively rare, was the primary motivation behind this paper. By constructing eleven asset indices and three income measures, this paper has explored how household rankings and the estimated impact of resource availability on child nutritional status changes when different constructs are used.

A number of related points, on the implications of using different measures of economic status, have emerged:

First, income measures tend to be highly correlated, regardless of whether they adjust for household size and spatial variation in prices or not. Also, compared to price-adjusted *per capita* income measures, failure to adjust for price variation exerts only a small influence on the estimated effect of income on nutritional status, with the direction of the effect depending on the quintile. However, failure to take household size into account produces quintile coefficients that are very much smaller – by a magnitude of 0.1 to 0.2 of a standard deviation than in *per capita* income models. This suggests that adjusting for household size in models of child nutritional status is clearly essential, but collecting data that allows for spatial variation in prices may not be very important – at least for studies conducted within these seven provinces of China.

Second, asset indices are not highly correlated with each other, in general, but there is (unsurprisingly) some clustering within certain classes of indices, e.g. among the three indices constructed using PCA and among the three indices constructed using median monetary value data. While, on the whole, regressions using asset indices did not produce similar quintile coefficients for the different asset measures, the range of coefficients in the poorest quintile was very tight, ranging from 0.72 to 0.82. The universal significance and narrow range of coefficient values in the poorest quintile suggests that the rankings produced by the different asset measures are more consistent at the extremes of the income distribution than in the middle. The practical implication of this finding is that if the focus of research is on health disparities between rich and poor, or motivated predominantly by concern for the position of the poorest of the poor, then using almost any one of these asset measures will yield very similar results.

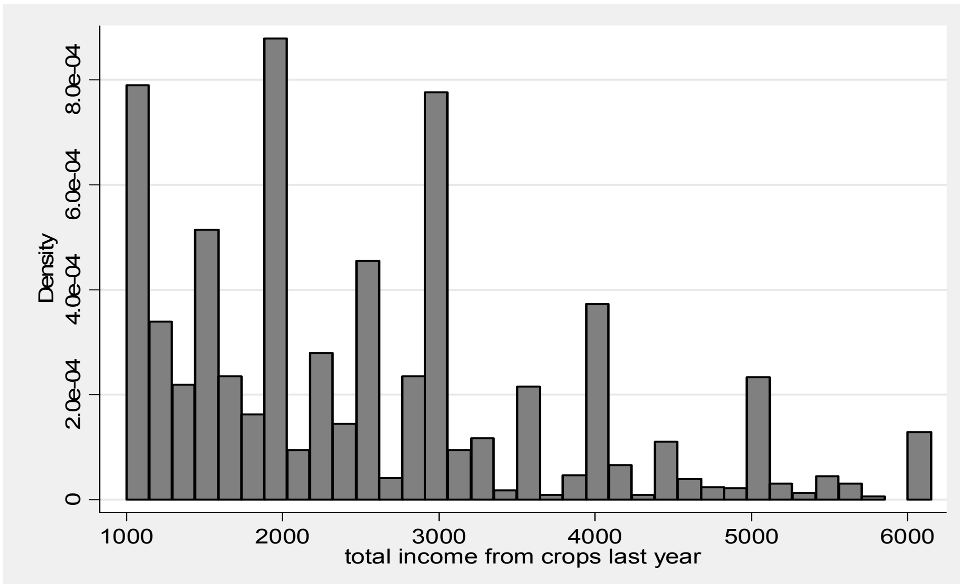
Third, as evidenced by the regression models, income measures appear to produce weaker relationships between economic status and child nutritional status than asset measures, or (stated differently) yield smaller child health disparities – at least in China, and at least for the set of assets examined in this paper. For example, in models using the *per capita* income measures, the height-for-age z-score of children in the first (poorest) quintile was between 0.63 and 0.67 standard deviations smaller than the height-for-age z-score of children in the upper quintile, but in models using asset indices it was between 0.72 and 0.82 standard deviations smaller.

Fourth, of all measures of economic status, only the *per capita* income measure produces a pattern whereby the magnitudes of the coefficients increase progressively, and significantly, with each successive quintile. In other words, it is the only construct that yields the expected result that, *ceteris paribus*, children in each successively lower quintile are in worse health status than children in the previous quintile. While no asset measure produces this relationship for all quintiles – a result which may in part reflect relatively small changes in asset ownership in the upper quintiles (compared to relatively large changes in income) – three asset measures do so for the three lowest quintiles: the simple asset index, the asset index that accounts for the number of assets, and the index constructed using dummy variable PCA.

Fifth, while the results unfortunately provide little guidance on whether to prefer an income or an asset construct in models of child nutritional status, they do make two clear points: (i) they highlight the fact that the choice of construct – income or assets – is far from inconsequential in terms of the magnitude and significance of the estimated coefficients, and (ii) they provide an indication of the potential direction and magnitude of bias if one class of measures is chosen over another. With respect to the latter, they caution that when one uses an income measure rather than an asset index, one risks underestimating the impact of extreme poverty on child nutritional status, but is more likely to observe an effect of economic status in the upper-middle quintiles. When using asset measures rather than income measures, on the other hand, one risks overestimating the impact of poverty on child nutritional status and is less likely to observe an effect of economic status in the upper-middle quintiles.

Appendix: Additional figures and tables for Paper II

Figure 17: Evidence of recall bias: lumping of data at intervals of 1000 yuan and 500 yuan



Source: China Health and Nutrition Survey, 2000

Note: This graph of household's self-reported earnings from the sale of crops during the previous year highlights the effect of recall bias. There is clear evidence of rounding at multiples of 1,000 yuan, with smaller spikes at intervals of 500 yuan.

Table 15: Weights assigned to each asset, equivalent to reciprocal of proportion of households owning the asset

Asset	Proportion of sample owning	Reciprocal/Weight
Original variables		
Radio	0.39	2.5
VCR	0.06	17.4
Black and white TV	0.35	2.8
Color TV	0.64	1.6
Washing machine	0.49	2.1
Refrigerator	0.38	2.6
Airconditioner	0.11	9.4
Sewing machine	0.45	2.2
Fan	0.84	1.2
Camera	0.12	8.5
Microwave oven	0.06	16.8
Rice cooker	0.40	2.5
Pressure cooker	0.46	2.2
Personal computer	0.05	21.7
Telephone	0.46	2.2
DVD player	0.25	3.9
Tricycle	0.12	8.6
Bicycle	0.68	1.5
Motorbike	0.21	4.7
Car	0.03	33
Constructed variables		
VCR or DVD player	0.27	3.7
Any TV	0.90	1.1

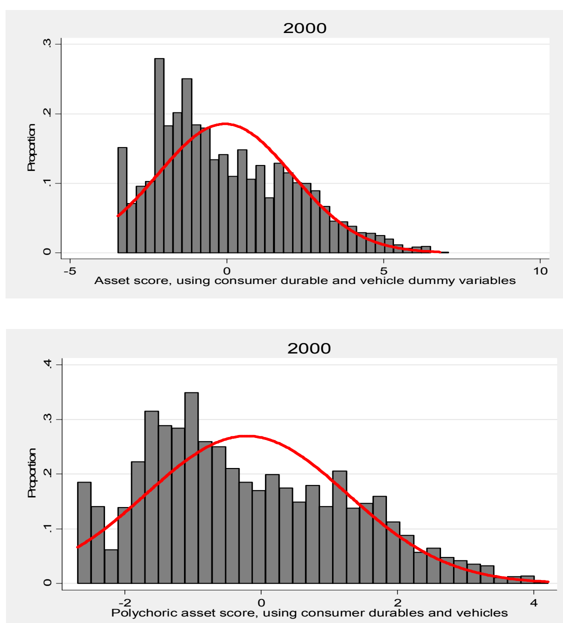
Source: China Health and Nutrition Survey, 2000

Table 16: Median price of asset variables in sample and by province

	Sample	Jiangsu	Shandong	Henan	Hubei	Hunan	Guangxi	Guizhou
Radio	110	150	200	100	100	150	100	115
VCR	1000	1000	1800	2000	1400	900	700	800
B/W TV	200	200	300	200	200	200	250	200
Color TV	1800	2000	2100	1900	1750	2000	1600	1500
Washing machine	500	500	700	500	500	500	900	500
Fridge	1500	1300	2000	1800	1200	1400	1500	1200
Aircon	3300	4000	6000	2300	3000	3000	3700	2100
Sewing machine	140	150	195	120	110	150	150	128
Fan	150	200	200	150	150	180	150	100
Computer	3000	4500	5000	2000	5000	960	1200	1000
Camera	300	400	300	250	300	290	500	235
Microwave	700	615	800	500	775	750	800	500
Rice cooker	100	100	180	120	100	100	80	80
Pressure cooker	80	100	100	93	80	80	70	80
Telephone	600	1200	600	555	600	600	500	200
DVD player	1000	1000	1000	1000	800	800	800	800

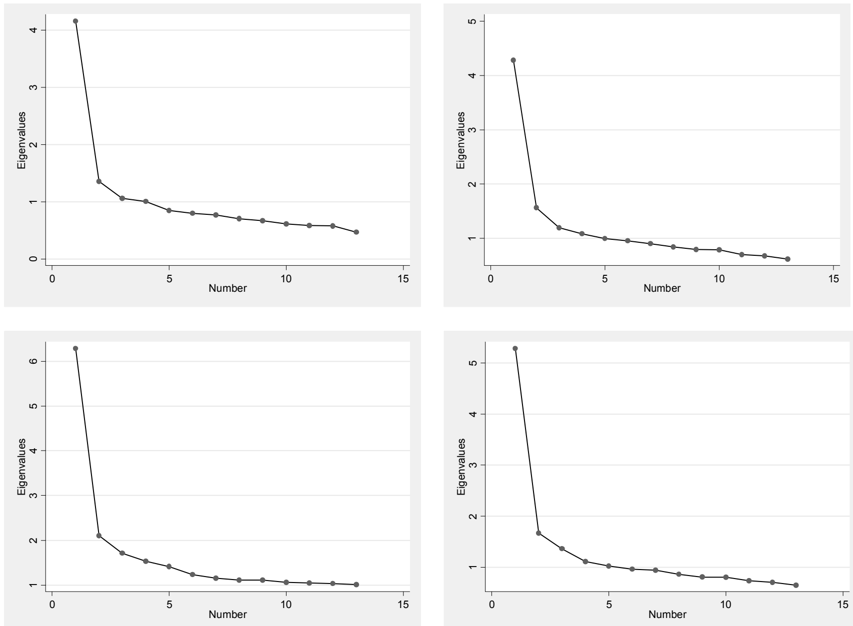
Source: China Health and Nutrition Survey, 2000

Figure 18: Wealth distribution constructed by running ordinal and polychoric PCA on dummy consumer durables and vehicles



Source: China Health and Nutrition Survey, 2000
Note: Assets include consumer durables and vehicles only

Figure 19: Scree plots



Source: China Health and Nutrition Survey, 2000
Note: Graphs only display a maximum of 13 eigenvalues/ principal components; top left is for dummy variable PCA with consumer durables only; top right is for dummy variables PCA with consumer durables and vehicles; bottom left is for ordinal PCA with consumer durables, vehicles and dummy-coded utility variables; bottom right is for ordinal PCA with consumer durables, vehicles and ordinally-coded utility variables.

Table 17: Internal coherence of PCA-constructed asset indices, i.e. percentage ownership of select assets across asset tertiles

	Dummy utilities			Ordinal utilities			Polychoric with ordinal utilities		
	T1	*T2*	*T3*	*T1*	*T2*	*T3*	*T1*	*T2*	*T3*
Car	0.64	3.04	5.16	0.46	3.04	5.35	0.46	3.04	5.35
Bicycle	58.38	70.41	72.26	53.82	72.51	74.72	53.55	72.51	75.00
Television	72.28	96.87	99.63	70.78	98.15	99.82	70.60	98.34	99.82
Pressure cooker	13.35	47.74	76.31	12.81	47.23	77.12	12.90	47.05	77.21
Fridge	1.57	24.33	85.81	1.01	22.32	88.1	0.83	22.79	87.82
Washing machine	8.38	43.87	91.06	6.54	44.10	92.53	6.51	44.00	92.62
Fan	64.92	89.86	95.86	63.13	90.96	96.49	62.76	91.33	96.49
In-house flush toilet	1.1	23.59	77.7	2.58	27.24	72.42	3.04	27.42	71.77
In-house tap	14.92	51.34	90.6	21.75	49.63	85.33	22.21	49.82	85.42

Source: China Health and Nutrition Survey, 2000
Note: T is for tertile, where T1 is the poorest tertile and T3 is the wealthiest tertile

Table 18: Internal coherence of three select asset indices, i.e. percentage ownership of select assets across asset tertiles

	Researcher assigned weights			Self-reported monetary value			Sample median prices		
	T1	*T2*	*T3*	*T1*	*T2*	*T3*	*T1*	*T2*	*T3*
Car	0.40	3.22	5.60	0	0.12	6.01	0.93	4.07	4.48
Bicycle	49.52	76.40	79.18	60.23	72.33	79.54	63.64	70.73	74.23
Television	73.92	98.86	99.81	78.39	98.84	100.00	83.24	99.41	99.81
Pressure cooker	16.40	50.73	75.54	21.84	44.42	67.28	22.01	49.75	71.32
Fridge	2.72	26.51	87.02	2.53	30.23	77.34	0.55	20.80	95.22
Washing machine	9.12	48.54	92.16	11.84	50.00	81.73	12.34	45.73	91.87
Fan	64.40	93.35	97.29	73.10	89.07	96.30	77.16	89.99	95.22
In-house flush toilet	10.51	28.11	66.95	9.24	30.19	57.13	8.89	31.37	65.52
In-house tap	31.50	47.65	80.37	30.29	52.87	68.48	29.68	50.49	78.33

Source: China Health and Nutrition Survey, 2000
Note: T is for tertile, where T1 is the poorest tertile and T3 is the wealthiest tertile; these indices did not include utilities in their construction

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PAPER III

CHILD NUTRITIONAL STATUS IN CHINA IN THE 1990S:

DID INCOME MATTER?

1. Introduction

During the 1990s, the prevalence of child undernutrition in China was halved – the largest decline in that decade in the world. Between 1991 and 2000, the national prevalence of underweight fell by 41% and the prevalence of stunting by 56% (calculated from WHO 2004). As is well known, the 1990s was also a period of dramatic and sustained economic growth, averaging around 9.7% *per annum* from 1990 to 1999 (calculated from Morrison 2005) which, for many, translated into substantial income gains. Since income is a key determinant of a household's decision to investment in its children's nutritional status (see, for example, the reviews by Behrman and Deolalikar 1988; Strauss and Thomas 1995; Chamarbagwala *et al.* 2004), this raises the question as to whether, and to what extent, variation in nutritional status in China during the 1990s can be attributed to variation in household income.

The pathway from income to nutritional status is a complex interaction of immediate, underlying and basic determinants (Mosley and Chen 1984; UNICEF 1990; Jonsson 1993; Smith and Haddad 2000). Other socioeconomic factors may mitigate or accentuate the income effect, and for any one country, community or household the exact magnitude of the effect of income on child nutritional status will depend on the specific constellation of factors at play. In light of the broad range of policies and interventions that can affect health status, it is not surprising that cross-country studies of the income elasticities of child nutritional status yield wide-ranging estimates. For example, a recent twelve-country comparison obtains income elasticities of malnutrition (as captured by height-for-age z-scores) that vary from 0.039 in Pakistan to 0.44 in Morocco (Haddad *et al.* 2003).

In this paper, the focus will be on the role of income as a determinant of child nutritional status in China and how its effect has been mediated by two specific areas of policy. First is the one child family policy that has been associated with a sharp reduction in fertility rates – the total fertility rate fell from 2.8 when the policy was first implemented in 1979 to 1.8 in 2000 (Festini and de Martino 2000). Second is a protracted period of health system reform which has resulted in increased variation in the accessibility, cost and quality of healthcare, with potential implications for children's consumption of health services, their incidence of illness and, thus, nutritional status.

The paper commences with an elaboration of the theoretical model and a discussion of the data, and then provides a brief empirical description of the trends in child nutritional status in China during the 1990s. Then, it examines changes in income, the one child policy and healthcare access during this period. Finally, the empirical relationship between income and child nutritional status, and the extent to which that relationship is mediated by access to healthcare and being an only-child is investigated using ordinary least squares (OLS), random effects (RE), fixed effects (FE) and instrumental variables (IV) models. Data are drawn from four waves of the longitudinal China Health and Nutrition Survey (CHNS).

2. Theoretical framework

The theoretical approach adopted in this paper has its origins in Becker's microeconomic model of household production (Becker 1965; Becker 1981) and the expansion of this model to the demand for health (Grossman 1972). The application of this model to child health status is well-documented (Behrman and Deolalikar 1988; Strauss and Thomas 1995; Currie 2000).

It assumes a one-period unitary model of decision-making whereby a household maximizes a **utility function** that depends on the consumption of commodities including a composite (market-purchased) household consumption good (X_{jt}), the consumption of leisure by all household members (L_{jt}) and child health or nutritional status (H_{ijt}), conditional on a set of household taste and preferences shifters (W_{jt}):

$$U_{jt} = U(X_{jt}, L_{jt}, H_{ijt}; W_{jt}) \tag{1}$$

The household faces three constraints:

(i) a health production function representing the technology available to the household to transform available inputs into the health of the child. The health or nutritional status of child *i* in household *j* in period *t* can be described by the function

$$\mathrm{H}_{\mathrm{ijt}} = h(M_{ijt}, Z_{ijt}, V_{ijt}; G_{ijt}, F_{jt}) \tag{2}$$

where nutritional status depends on M_{ijt}, the endogenous material health inputs (such as immunizations, nutritional intake, healthcare utilization); on Z_{ijt}, the exogenous characteristics of the child and the household in which it lives; and on $\mathrm{V}_{\mathrm{ijt}}$, the endogenous child health-related time inputs. Child health is also conditional on two vectors of exogenous child-specific (G_{ijt}) and household/community-specific (F_{jt}) variables that are time-varying or time-invariant indicators of health technology[53]. As such, they influence the choice of health inputs and/or the efficiency with which existing inputs are combined to produce child health, and in so doing affect the shape of the health production function. Exogenous child-specific factors (G_{ijt}) could include age and gender. An example of an exogenous endowment at the household-level (F_{jt}) is the education of the caregiver (typically the mother) who makes health-related decisions that impact on the child's health. At the community-level, prevailing cultural norms regarding child health practices, the availability of healthcare facilities, relative food prices and communication or transportation infrastructure are examples of F_{jt} variables.

(ii) a **time endowment** reflecting the total time available to the household (N_{jt}) to allocate between wage labor (T_{jt}), leisure (L_{jt}) and health-related activities (V_{jt})

$$N_{jt} = T_{jt} + L_{jt} + V_{jt} \tag{3}$$

(iii) the **budget constraint** representing the total financial resources, consisting of wage income ($w_t T_{jt}$), non-wage income (Y_{jt}) and net assets (A_{jt}), available to the household with which to purchase market-produced health-related inputs (M_{ijt}) and other consumption goods (X_{jt}) at prices p^M and p^X respectively

$$p_t^X X_{jt} + p_t^M M_{ijt} = w_t T_{jt} + Y_{jt} + A_{jt} \tag{4}$$

[53] F_{ijt} and G_{ijt} variables can be thought of as exogenous productivity shifters, which may be either permanent or temporary shocks (Currie 2000).

Solution of the household's optimization problem yields the reduced-form health demand equation[54]:

$$\mathrm{H}_{\mathrm{ijt}} = h(m(p_t^X, p_t^M, w_t, N_{jt}, Y_{jt}, A_{jt}; G_{ijt}, F_{jt}), Z_{ijt}, v(p_t^X, p_t^M, w_t, N_{jt}, Y_{jt}, A_{jt}; G_{ijt}, F_{jt}); G_{ijt}, F_{jt}) \tag{5}$$

The particular functional form *h*( ) depends on the underlying functions characterizing household preferences and health production.

In reduced-form health demand models, all exogenous prices and endowments enter into the determination of each of the endogenous variables, implying that the prices of *all* goods (and not only the prices of health-related input prices) and all individual and household endowments have an effect on the health of children in the household.

3. Data

Data are drawn from a large-scale panel survey conducted in nine provinces of China from 1989 to 2004, the China Health and Nutritional Survey (CHNS). This paper utilizes data from four of the six waves, from 1991 through 2000[55] and is restricted to the seven provinces of Jiangsu, Shandong, Henan, Hubei, Hunan, Guangxi and Guizhou[56].

In each of the provinces, a multistage, random cluster procedure was used to draw the sample. Counties in each province were stratified by income (low, middle, and high) and a weighted sampling scheme was used to randomly select 4 counties in each province. In addition, the provincial capital and a lower income city were selected where feasible. Villages and townships within the counties, and urban and suburban neighborhoods within the cities, were selected randomly.

The first CHNS wave, in 1989, surveyed 15,917 individuals. In CHNS 1991, only individuals belonging to the original sample were interviewed, resulting in a total of 14,778 individuals. In CHNS 1993, all the new households that had been formed from the 1991 sample households who resided in sample areas were added, resulting in a total of 13,893 individuals. In CHNS 1997, all newly-formed households who resided

[54] The derivation of this health demand equation is given in the Appendix.
[55] In 1989, anthropometric data were not collected for children over the age of six years, and 2004 data have not yet been released.

[56] The provinces of Liaoning and Heilongjiang are excluded because the former lacks data for 1997 and the latter lacks data for 1989, 1991 and 1993.

in sample areas were once again included, plus additional households and communities to replace those no longer participating. Also, Heilongjiang province replaced Liaoning province, bringing the sample size to 14,426 individuals in 1997. In CHNS 2000, newly-formed households, replacement households, and replacement communities were again added, and Liaoning province returned to the study. In 2000, the sample consisted of 15,648 individuals. The total number of primary sampling units increased from 190 in 1991 and 1993 to 216 in 2000[57].

In panel studies, there is always the concern that attrition will bias the sample in a systematic fashion. In particular, if the parameters governing the missing data process are likely to be related to the parameters that need to be estimated, then this "missingness" is not of an ignorable nature. There has clearly been some attrition in this sample: of the 3,207 households included in 1991 CHNS and residing in the seven provinces examined in this paper, 92% were in the 1993 sample, 82% were in the 1997 sample and 76% were in the 2000 sample (own calculations). However, there is no *a priori* reason to expect that particular types of households are more likely to have left the CHNS study than others. Moreover, in each wave of data collection, careful attention was given to the selection of replacement households so that the number of households in each cross-section remained approximately constant across the decade, with the exception of a small dip in 1991[58].

4. Nutritional status: measurement and evidence from the 1990s

The dependent variable, nutritional status, is captured by age- and sex-standardized height-for-age z-scores[59], constructed on the basis of the 2000 CDC growth charts. Children who have a height (or

[57] These figures refer to all nine province included in the CHNS study. In the sample of seven provinces and four waves that are analyzed in this paper, there were 13,229 individuals in 1991, 12,492 individuals in 1993, 14,268 individuals in 1997 and 13,657 individuals in 2000.

[58] There were 3,207 households in the seven provinces in 1991, 2,986 households in 1993, 3,296 households in 1997 and 3,237 households in 2000.

[59] Since children under the age of 36 months were measured using length boards, recumbent length-for-age rather than height-for-age is used for children in this age group.

recumbent length)-for-age measurement that is less than two standard deviations below the median value of the NCHS/WHO reference group are referred to as stunted[60].

There is convincing evidence that the NCHS/WHO standards are an appropriate yardstick by which to measure the growth of Chinese children. Cross-national studies have shown that, once allowances are made for socioeconomic strata (Habicht *et al.* 1974) and variations in maternal height and feeding patterns (WHO 2000), children from different parts of the world tend to exhibit remarkably similar growth trajectories. Also, studies of ethnic Chinese children born in developed countries indicate that they tend to follow similar growth patterns to the NCHS reference population (Toselli 2005; Sit *et al*. 2001), suggesting that it is more likely to be environmental rather than genetic factors that contribute to any growth retardation observed among children resident in China. Third, the use of the NCHS reference avoids a major disadvantage of relying on country-specific growth standards, namely that any systematic bias or discrimination against specific subgroups, such as girls, is hidden (Harriss-White 1997).

Following the recommended exclusion ranges for implausible values, height-for-age z-scores that are less than -5 and greater than +3 have been excluded from the analysis (WHO 1995). After pooling data from seven provinces across four years, there are more than 7,500 anthropometric observations (specifically, 7,587 height-for-age data points) for children under the age of twelve (i.e. under 144 months). In this age group, the sample contains 2,441 height-for-age observations in 1991, 2,161 children in 1993, 1,726 children in 1997 and 1,259 in 2000.

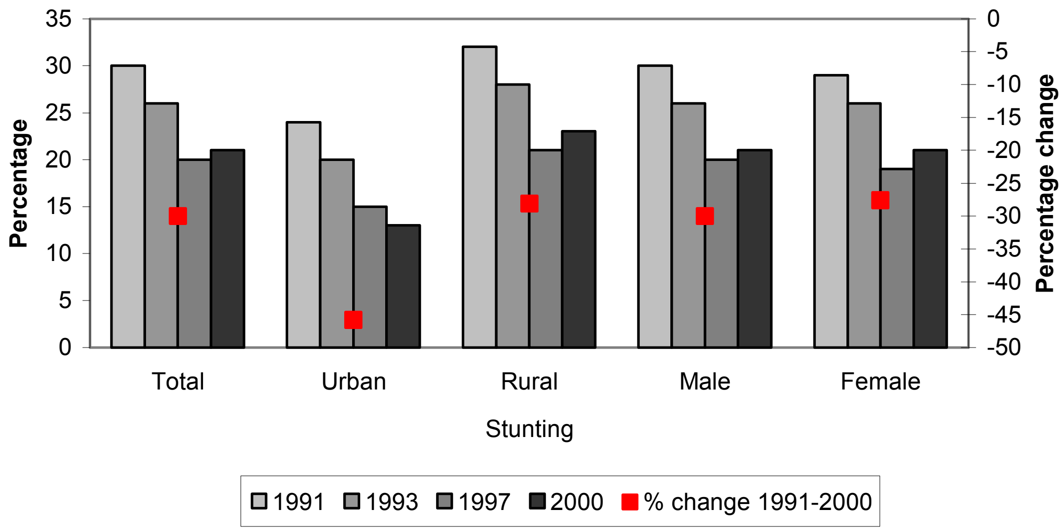
Stunting indicators reveal an unambiguous, and large, reduction in the prevalence of child undernutrition in the CHNS sample during the 1990s (see Figure 1)[61]. From 1991 to 2000, the prevalence of undernutrition fell by about one third, although the change was greater in the early and middle period than towards the end of the decade. Also, the prevalence of stunting fell by a much greater percentage in urban (46%) than in rural areas (28%). Despite these dramatic improvements, in 2000 there remained a sizeable percentage of children that suffered from stunting (21%). This percentage was much greater in rural than in

[60] The term "severely stunted" is used when the measurement is less than three standard deviations below the reference median. "Mildly stunted" refers to a measurement of less than one standard deviation below the reference median.

[61] Since figures are not age-adjusted and the proportion of older children – who are more likely to be malnourished (Shrimpton *et al.* 2001) – increases relative to younger children with each subsequent wave, these figures may even understate the full extent of the reduction in the prevalence of undernutrition.

urban areas, and among older children than among younger children, but there does not appear to have been a marked gender differential.

Figure 20: Percentage change in the prevalence of stunting, by subgroup, 1991-2000



Source: China Health and Nutrition Survey, 1991-2000

5. Changes in income during the 1990s

The 1990s were years of remarkable economic growth in China that, still today, shows little sign of slowing. The literature suggests that this economic growth appears to have been accompanied by considerable household income growth, at least among some segments of the population.

In rural areas, Ravallion and Chen (2004)[62] found that mean household *per capita* income in 1980 prices rose from 360 yuan in 1991 to 617 yuan in 2000. Using the Government of China's poverty line[63], the United Nations (2004) calculates that the percentage of the rural population living in poverty fell from 9.4% to 3.2% between 1990 and 2000. Agricultural growth appears to have been the key explanatory factor (World Bank 2001). Progress was uneven over time, though: half of the decline in rural poverty came in the first few years of the 1980s, with poverty reduction stalling during the late 1980s and early 1990s,

[62] This study uses National Bureau of Statistics (NBS) data from the Rural Household Surveys (RHS) and Urban Household Surveys (UHS).

[63] The official poverty line for rural areas was equivalent to 300 yuan per person at 1990 prices (Ravallion and Chen 2004).

recovering pace in the mid-1990s, and slowing again in the late 1990s (Ravallion and Chen 2004). Indeed, Benjamin *et al.* (2005), using a different rural dataset[64], found that the absolute living standards of the poor declined so much from 1995 to 1999 that they approached the income levels of 1987 and that as many as a half of these households were not unambiguously better off in 1999 than in 1987.

Living standards in urban China have risen more sharply than in rural areas. Ravallion and Chen (2004) found that national mean household *per capita* income at 1980 prices rose from 478 yuan in 1991 to 916 yuan in 2000[65]. This means that urban (income) poverty has effectively been eliminated (Benjamin *et al.* forthcoming), in contrast to rural China where the income gains of the first half of the 1990s deteriorated.

In both urban and rural areas, inequality has been increasing, but more so in urban areas so that over time, there has been some convergence in the levels of inequality in urban and rural areas (Benjamin *et al.* 2005; Démurger *et al.* 2005[66]). The national Gini-coefficient increased from 0.342 to 0.407 between 1989 and 2000 (Wu and Perloff 2004).

In the CHNS sample, somewhat different income trends are observed (see Figure 21). In the sample, the mean real *per capita* cost-of-living adjusted income has, in fact, fallen slightly over the course of the decade – by 14.8% in the total sample, by 9.3% in urban areas and by 18.1% in rural areas. Increases in real *per capita* income were observed over the period in Jiangsu and Guangxi provinces, however (see Appendix, Table 24 for exact figures). One possible explanation for the difference between results obtained using the CHNS data and results obtained in other studies could be the very different samples. The provinces included in the CHNS are a small fraction of the total 22 Chinese provinces[67]. Another reason may be different approaches taken to the measurement of income, especially the range of income sources

[64] The Research Centre for Rural Economy (RCRE) has collected data in nine provinces (namely Jilin, Shanxi, Henan, Hunan, Anhui, Jiangsu, Guangdong, Sichuan and Gansu), including 100 villages and 7,000 to 8,000 households every two years since 1987. The data used by Benjamin *et al*. (2005) cover rural areas from 1987 to 1999.

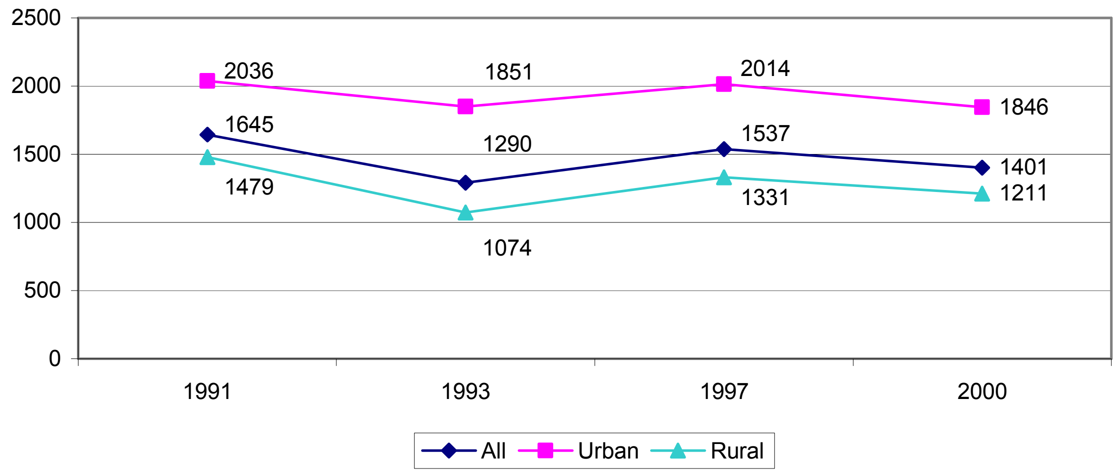
[65] When adjustments are made for urban-rural cost-of-living differentials, these figures become 617 yuan in 1991 and 1,015 yuan in 2000.

[66] This study use data from the three waves (1988, 1995 and 2002) of the China Household Income Project (CHIP) which covers 10 provinces (plus Sichuan in the 1995 wave), and is a sub-sample of the sample selected by the NBS in its annual household survey. It focuses on urban areas.

[67] Officially, there are 34 province-level administrative units in China including 22 provinces, five autonomous regions, two special administrative districts (Hong Kong and Macao), four directly-administered municipalities (such as Beijing and Shanghai) and Taiwan.

included. The income measure used in this paper aggregates multiple net cash income streams, the market value of in-kind income, and the market value of home production consumed. Since, prior to 2000, no official consumer price index existed for China, another factor may be differences in the adjustments that each study makes for geographic and temporal price differences. A unique consumer price index that was developed for use with the CHNS data is used in this paper, and its construction is discussed in more detail in the Appendix.

Figure 21: Trends in mean real *per capita* income, by urban and rural area in the seven provinces, 1991 to 2000, in real 1989 yuan



Source: China Health and Nutrition Survey, 1991-2000
Note: Means are calculated at constant 1988 yuan

6. The one child policy and health system reform: implications for nutritional status

<u>6.1 The one child policy</u>

The one child policy is one of the more recent of a cornucopia of population control measures that have been in effect in China since at least as early as the 1950s[68]. Introduced formally in January 1979, it

[68] Measures enacted over the years include delaying the age of marriage, extending the interval between marriage and birth, lengthening the space between births, and near-compulsory contraceptive use, especially after the birth of the first child (Scharping 2003).

provided incentives to couples who signed the "one child certificate" and discouraged, and later penalized, second births. The one child certificate offered rewards such as extra food rations, better housing, health subsidies and free education to the single child born in urban areas, while in rural areas, increases in the allotment of farmland was common. Punitive measures included fines, the loss of jobs or the denial of workplace promotions or privileges. As the policy tightened, second births were forbidden altogether, except in extraordinary circumstances. The strictness of the policy reached its peak around 1983 before the release of Central Document 7 which condemned coercive enforcement and initiated an expansion of the range of conditions under which couples could have two children (Li 2004). By 1988, about 12% of the Chinese population lived in areas with global second-child permission and a further 50% lived in provinces extending a second child permit to peasant households with only one daughter (Scharping 2003). Other concessions included families where there was a second marriage and the one partner had not had a child, where the first child had an abnormality, families where the father had a dangerous occupation and when both spouses were only-children. As the 1990s progressed, there was a tendency to reject further extensions of second-child permits, though. Scharping (2003) estimates that, in 2003, about 60% of the Chinese peasant population and about 5% of the urban population was eligible for second-child permits. While the policy is considerably more relaxed in rural than in urban areas, there is substantial variation in its rules and conditions, as well as in the stringency of its application – even across fairly small geographic areas – and these have fluctuated over time[69].

In this analysis, the effect of the one child policy on child nutritional status in China is not captured directly, but rather through its impact on whether or not the index child is an only-child[70]. Given China's

[69] Throughout its period of its implementation, the one child policy has been a very much decentralized policy: no national birth-planning regulations for the whole of China were ever passed; most of the details of implementation were delegated to provinces; and, provinces granted much latitude to subordinate levels, such as prefectures and municipal regions, cities, counties, villages and work units, in the interpretation of existing policies and the initiation of new policies.

[70] While the community-level measures of the severity of the one child regulations contained in the CHNS appear attractive because of their exogeneity, they suffer from two persuasive drawbacks. First, since the number of siblings in a household in the current period is determined by the one child policy stipulations in previous periods, data on one child policy regulations from the 1980s would be needed to explain health outcomes for children in the CHNS sample (1991 to 2000). Second, even if one were to make an argument that current period regulations are a sufficient representation of past regulations (perhaps by arguing that spatial variation in policy is greater than temporal variation, for example), few of these policy variables are available in all years of data collection and those that are have many missing values.

one child policy, it is expected that this variable will affect child nutritional status in a multitude of ways, many of which operate through the income variable.

First, having one child rather than many children increases the *per capita* availability of household resources that can be directed towards enhancing child nutritional status. Second, it may also increase the *total* household resource availability since women who have fewer children may spend more time in wage-employment (Entwisle and Chen 2002). Third, the incentives and sanctions associated with the one child policy have implications for child health and nutritional status: some may act through an income effect (e.g. fines may reduce money available for child-raising or grants for single children may increase it)[71] and others through a price effect (e.g. where healthcare subsidies are provided for only-children). Fourth, to the extent that the one child policy may alter prenatal and obstetric care-seeking behavior, it holds implications for birth outcomes which, in turn, are correlated with nutritional status later in childhood. Doherty *et al.* (2001) show that the one child policy has been a significant deterrent to the utilization of prenatal and obstetric care among women with unapproved pregnancies. On the other hand, it has also been argued that the close monitoring of compliance with some of the terms of the one child policy, such as compulsory contraception, may have brought many more women into contact with healthcare services than would otherwise have been the case (Festini and de Martino 2004).

For the last two reasons, it is expected that the magnitude of the effect on nutritional status of being an only-child will be greater in China than is typically found in other countries where the effect of being an only-child operates only through the first two mechanisms.

CHNS data highlight the trends in the percentage of single children. Although not all spatial and temporal differences in the percentage of only-children can necessarily be attributed to the one child policy, one plausible measure of the effect of the one child policy – the percentage of children under twelve without siblings (under the age of 18) – indicates an overall upward trend in the number of only-children[72]. Between 1991 and 2000, the percentage of children without siblings rose from 27.1% to 42.8%, with the

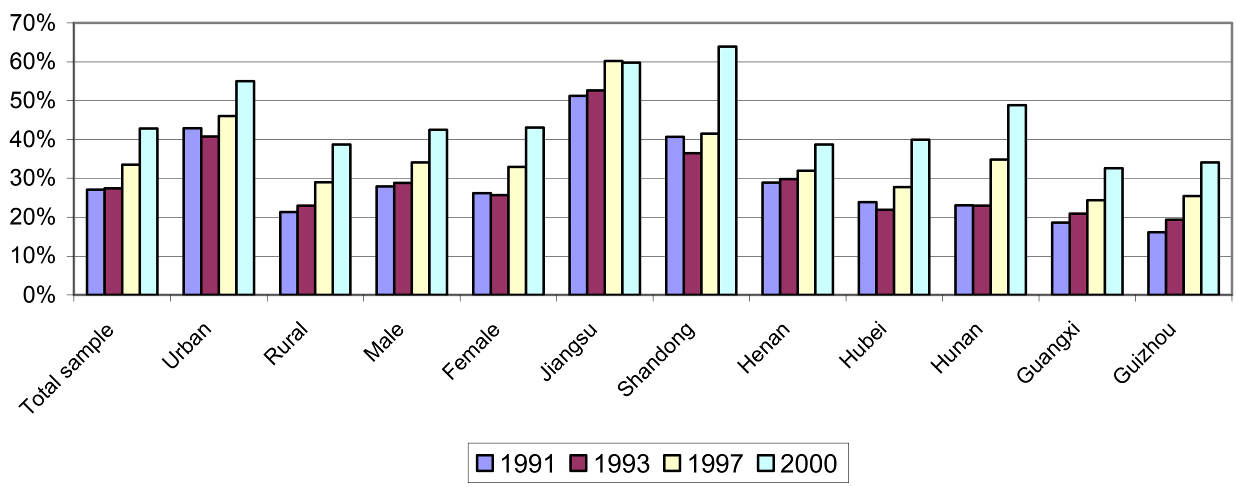
[71] The income effect will be captured by the "only-child" variable since this income source is not included in the constructed income measure.

[72] Another indicator is that the mean age of children under 12 migrates upwards in each successive survey year, increasing from 6.36 in 1991 to 7.37 in 2000.

sharpest changes occurring in the latter half of the decade, as would be expected in light of the tightening of the policy about ten years earlier around the time of the children's birth.

Spatial variation, across urban-rural areas and across provinces, is marked. The percentage of urban children that are only-children is consistently higher than the percentage of rural children that are only-children, but the gap has narrowed over time with the percentage of only-children in urban areas increasing by 28.4% between 1991 and 2000, and the percentage of only-children in rural areas increasing by 81.7%. In general, it is the more economically developed provinces (such as Jiangsu and Shandong) that have a higher percentage of only-children, but there too, the gap is narrowing with the largest changes occurring in those provinces that entered the 1990s with a smaller percentage of only-children (such as Guizhou). Once again, while these differences cannot be conclusively attributed to the one child policy, the spatial variation in the percentage of only-children does approximately correspond with the urban-rural and inter-provincial severity of the one child policy regulations[73].

Figure 22: Percentage of children, under 12, without any siblings, 1991 to 2000



Source: China Health and Nutrition Survey, 1991-2000
Notes: For each category of disaggregation, the percentage of children without any siblings in 2000 is statistically different from, and larger than, the percentage of children without any siblings in 1991. Exact figures are shown in the Appendix, Table 25.

[73] Scharping (2003) provides an exceptionally detailed account of the timing of the implementation of different one child policy regulations and incentives, disaggregated by urban-rural areas within provinces.

6.2 Health system reform as a determinant of nutritional status

Along with income, the provision of health services is a potentially important determinant of child nutritional status, and the role of economic development versus health services in improving the health status of populations is a long-standing, and unresolved, historical debate in developed (see, for example, McKeown *et al.* 1975, Preston and Van de Walle 1978) and developing countries (see,for example, Caldwell 1986; Glewwe *et al.* 2002; Banister and Zhang 2004).

Reducing the incidence, severity or duration of childhood disease is one clear theoretical pathway by which access to preventive and curative healthcare services can be expected to positively influence nutritional status. Children, and especially young children, who fall ill may rapidly deplete their nutritional stores because of reduced intake of food, poor absorption of nutrients and the increased demands of combating disease (Scrimshaw and SanGiovanni 1997; Allen and Gillespie 2001), leading to growth retardation. A second theoretical pathway is via the provision of appropriate antenatal care and advice to pregnant women which may improve birth outcomes and nutritional status later in life (Schmidt *et al.* 2002). Thus, to the extent that health facilities can be accessed for quality preventive and curative services, their presence should be associated with improved nutritional status. Indeed, the presence of a public hospital in the municipality has been found to have a positive effect on child nutritional status in Colombia, as has the expansion of health infrastructure in Peru (Behrman and Skoufias 2004).

In China, contemporaneous with market-orientated reforms in other sectors of the economy, the health sector has undergone substantial liberalization and change. In addition to the emergence of private providers, the reform process has been characterized by two features.

First, there have been huge cut-backs in government financing of healthcare services – from full financing in 1980 to a fixed budget that covered only around 30% of hospital expenditures in the late 1990s (Hsiao 1995; Liu and Mills 2002). Cut-backs were especially severe for public health and preventive healthcare services (Liu 2004) and institutions at the lower levels of service delivery, such as township facilities, village clinics, and maternal and child health centers (Liu and Mills 2002). Combined with the fact that any losses had to be assumed by the institution, a new system of staff bonuses whereby bonuses had to be paid from institution profits, and the fact that little additional funding was available from local governments (Hsiao 2004), health institutions began to charge higher and higher user fees.

Second, there has been a reduction in the percentage of the population covered by health insurance schemes. In rural areas, the collapse of the rural Community Medical Scheme (RCMS) left many of those in rural areas without insurance coverage. Coverage of the rural population fell from 80-90% in 1979 to a low of 5.4% in 1985 before recovering slightly to around 10% in 1995 (Hsiao 1995). Many of the working urban population and their dependants are provided with health insurance through the *gongfei yilao* or Government Employee Health Insurance (GHI) and the *laodong baoxian* or Labor Health Insurance (LHI), but there too, coverage has fallen and members pay an increasing share of their medical expenses out-of-pocket. In the late 1990s, various experimental community-based health insurance schemes emerged (Akin *et al.* 2004) and a universal health insurance program was rolled out in all major cities (Wu *et al*. 2005), but information on the current extent of their coverage is not available.

Taken together, these two features mean that while cut-backs in government financing necessitated the increasing use of user fees, fewer Chinese were able to rely on health insurance schemes to help cover these costs. It is hypothesized that this reform process may have been detrimental to child health for a number of reasons:

First, as private medical providers emerged, the supply of healthcare facilities increased, but since private providers are poorly regulated, there is tremendous variation in the quality of care that they provide. Where there have been quality improvements, these have been accompanied by prices increases, so that quality improvements in the system tend to benefit those who can afford to pay for them (Liu 2004). There is some evidence that, on balance, public care is preferred to private care, although private clinics have been praised as being more conveniently located, having more flexible hours and being more responsive. "Consumers are forced to choose between technically better but costlier public health care, on the one hand, and cheaper private health care of questionable quality, on the other" (Lim *et al.* 2004: 232).

The second major consequence of the health reform process is that user fees may have become an important – and perhaps increasingly important – determinant of access to care. This is especially so in poor rural communities where most people are not employed by government or large firms and, so, are not covered by LHI and GHI insurance. Indeed, several studies have found that the cost of seeking treatment is a major barrier to care in China (see, for example, Lim *et al.* 2004), arguing that user fees have increased the delay in seeking treatment (Liu and Mills 2002), reduced the utilization of pre- and post-natal care[74] (Anson 2004), and contributed to the fall in child immunization coverage rates[75] (Liu and Mills 2002).

Third, there has been a shift in the nature of healthcare services provided – away from preventive and public health interventions and towards curative care – and a tendency to over-prescribe diagnostics tests and pharmaceuticals[76], all of which are likely to have been detrimental to investments in child health[77]. Since healthcare facilities are required to generate most of their own funding and the prices of many essential medical services are kept low (and sometimes below cost) through regulation, there is an incentive to over-prescribe so as to generate profits or even merely cross-subsidize the costs of essential services (World Bank 1997). Many preventive services are no longer financed by the government and healthcare providers receive little or no remuneration for providing preventive services (Liu 2004). Consequently, these services are often under-provided (Wong *et al.* 1995; Liu and Mills 2002).

These trends in healthcare provision are supported by the CHNS data. Between 1991 and 2000, the mean number of healthcare facilities that a household had access to increased slightly, but significantly, from 1.68 to 1.76. However, accessing them was not necessarily more convenient: for the facility located closest to the household, travel time did not fall between 1991 and 2000. If anything, access became more difficult with average travel costs increasing from a mean of 0.05 real yuan to 0.43 real yuan – although this may reflect increasing use of motorized transport, rather than bicycles or walking. The cost of care more than

[74] These price barriers persist despite the *Maternal and Infant Health Care Law of the People's Republic of China (MIHC Law) 1994* which committed the state to providing the "necessary conditions" and "material support" to make health services more accessible for mothers and children, particularly in remote and poor areas. The Law has been described as "largely advocatory rather than mandatory" (Tolhurst *et al.* 2004) and does not address the affordability of services and the variability of prices (Anson 2004).

[75] For example, it is argued that in Shandong province, immunization coverage fell from 75% in 1979 to 39% between 1984 and 1986 as a result of the introduction of user fees.

[76] It has been estimated that drug pricing practices inflate rural healthcare costs by at least once third (Hsiao 2004) and in Shanghai, it has been shown that between 1980 and 1989, the prescription fee was the fastest growing component of the outpatient consultation bill (Ho 1995).

[77] One relevant example of the unnecessary over-prescription of drugs is the treatment of child diarrhea which has come to involve unnecessarily long hospital stays, intravenous fluids and complicated drug treatments, rather than basic treatments, such as oral rehydration solution, which would be less costly (World Bank 1997).

doubled over the period: the typical (real) price of a cold or influenza-related visit increased from 3.8 yuan to 8.8 yuan.

To the extent that mean waiting time and the availability of medicines can be used as indicators of the quality of treatment available, there appear to have been some improvements. Mean waiting time to see a health care worker halved over the decade from a mean of 18 minutes down to 9 minutes, and while there was little change between 1991 and 1997, by 2000 there had been a statistically significant increase in the availability of medicine at facilities.

Table 19: Trends in accessibility and quality of health facilities in seven provinces, 1991 to 2000

	1991	1993	1997	2000
Mean no. of facilities available	1.68	1.71	1.68	1.76
For the closest health facility:				
Mean travel time (minutes)	10	9.7	9.8	10.7
Mean real cost of travel (real yuan)	0.03	0.06	0.07	0.13
Mean waiting time (minutes)	17.6	15.2	11.7	8.7
Mean real cost of treating a cold	3.8	5.6	6.2	8.8
% reporting medicine generally available	91.1%	93.3%	92.8%	93.6%

Source: China Health and Nutrition Survey, 1991-2000

Note: Information was collected on every facility that the household can use, and the means are calculated for the facility that is closest (in distance) to the household. Values are calculated for households with at least one child under 12. F tests show that the differences between mean values for 1991 and 2000 are all statistically significant at the 1% level, except for the "percentage reporting medicine generally available" which is significant at the 10% level.

The data also confirm that insurance coverage, which had declined towards the end of the 1980s (Henderson *et al.* 1995), continued to fall during the 1990s, but not by very much. The percentage of households that had an insurance plan fell from 26.6% in 1991 to 21.2% in 2000. Moreover, as the years progressed, fewer people's medical expenses were paid or partially paid by insurance. This reduction is evident across all types of care, but most pronounced for preventive care. There appears to have been some narrowing of socioeconomic differentials in insurance coverage: between 1989 and 1997, coverage doubled for the bottom quartile and fell for the top quartile[78] (Akin *et al.* 2004)

[78] This analysis did not include the 2000 wave of data.

Table 20: Trends in financial burden of treatment in seven provinces, 1991 to 2000

	1991	1993	1997	2000
% persons with insurance	26.6	21.7	23.9	21.2
% persons whose most recent curative healthcare was covered by insurance, if sought care in past month	78	66.4	67.1	55.1
% adults whose last preventive care treatment was at least partially covered by insurance, if sought care in past month	71.3	45.7	42.9	14.3
% child immunizations covered by insurance, if received immunization in past year	37.1	21.9	16.7	10.9
% women who report that their insurance covers prenatal care and delivery	51	36.1	21.1	18.1

Source: China Health and Nutrition Survey, 1991-2000

Note: F tests show that the differences for years 1991 and year 2000 are all statistically significant at the 1% level.

6.3 Other covariates of nutritional status – theory and measurement

In addition to income and policy variables, a number of other variables relevant to the production of child nutritional status are included in the model as covariates.

Child's age: Vulnerable children in developing countries tend to follow a growth pattern whereby the growth of infants starts to falter at about 3 months of age and declines rapidly until about 12 months, followed by a markedly slower decline until about 18-19 months (Shrimpton *et al.* 2001) and then stays constant or exhibits a slight catch-up. This pattern is linked to a number of factors including premature cessation of exclusive breastfeeding the failure to introduce appropriate complementary foods of sufficient quality and quantity between the ages of six and nine months, and exposure to disease (Allen and Gillespie 2001). To capture this pattern, z-scores are modeled as a quadratic function of age (in years).

Sex: Since data are already age- and sex-standardized, any observed effect of this variable cannot be attributed to biological differences in the growth patterns of boys and girls. Rather, differences are more likely to reflect social influences that result in differential health inputs, and hence health outcomes, for boys and girls.

Mother's education: The effect of mother's education on child nutritional status is a key variable of interest in a number of studies of the determinants of child nutritional status (see, for example, Thomas *et al* 1991; Glewwe 1999; Handa 1999; Christiaensen and Alderman 2004). One pathway along which mother's education or literacy can be hypothesized to affect child nutritional status is via its effect on household income (either through own income or through positive assortative mating), where higher incomes permit the purchase of more and better health-related inputs. Schooling is also thought to affect child health by improving women's access to information and enhancing their information-processing skills so that they can make better use of available health and nutrition inputs and more quickly adopt new health technologies and innovations. Another hypothesis, namely that mother's educational attainment may be proxying for the effect of genetic endowment on health status, is lent credence by Wolfe and Behrman's (1987) finding that maternal education is not significantly associated with child height once genetic endowment is controlled for. In this paper, mother's educational attainment is included in the model as a vector of four dummy variables. The first category is for mothers who have attained a maximum of 1 to 3 years of primary school, the second for those who have attained a maximum of 4 to 6 years of primary school, the third for those who have attained a maximum of 1-3 years of lower middle school, and the fourth is for those who have attained 1-3 years of upper middle school, 1-2 years of middle technical school[79] or education beyond secondary school (whether in a college or university setting). The omitted category is unschooled mothers. The dummy variable vector is preferred to a continuous measure of the number of years of schooling because each dummy represents a significant threshold of scholastic achievement.

Mother's height: Mother's height, measured in centimeters, is included to capture the child's unobserved growth potential or genetic endowment. In addition, mother's height has an effect on child height through its influence on birth weight (Xu *et al.* 1995; Blumenfeld *et al.* 2006). It may also be spuriously associated with the child's growth if it captures some unobserved household background characteristics that are health-promoting, and which affect both mother's height and child's growth. Father's height, which has been shown to be an important determinant of child birth weight and attained height (Magnus *et al.* 1984; To *et al.* 1998), is omitted from the model because of the large number of missing values on this variable, equivalent to 16.8% of the sample.

[79] In China, after completing the equivalent of lower middle school, children may attend 1 to 3 years of upper middle school or 1 to 2 years of middle technical school as part of secondary education.

7. Results

In the empirical formulation of the theoretical model outlined in section 2, the nutritional status of child i in community j at time t, (H_{ijt}), can be modeled as a function of k regressors that may be time-varying – such as the child's age, income, whether the index child is an only-child and access to health services – or time-invariant – such as the child's sex and genetic endowment.

Specifying a model with an individual-level time-invariant component to the error term, such as would be used if an individual fixed effects model were estimated, was explored, but abandoned. This is because estimating an individual level fixed effects model causes a number of time-invariant variables to drop out. Moreover, it is highly plausible that community-level unobservable factors, such as the general level of development in the community, transportation and communications infrastructure, may be a more important source of bias than individual-level unobservable factors.

Instead, this paper assumes that the composite error term ε_{ijt} has a random component μ_{it} – that varies across children and over time – and a fixed component v_j – that does not vary across children within specific communities over time, but may vary across communities in the sample. Year dummies capture any trending in nutritional status over time, i.e. aggregate time effects that have the same influence on the nutritional status of all children regardless of their other observed and unobserved characteristics.

$$H_{ijt} = \beta_0 + \beta_1 X_{ijt1} + \ldots + \beta_k X_{ijtk} + \delta_1 1993 + \delta_2 1997 + \delta_3 2000 + \varepsilon_{ijt} \, where \varepsilon_{ijt} = \mu_{it} + \nu_j \qquad (6)$$

The correct specification of a reduced-form child health demand model would not include a measure of household income. So, strictly speaking, this is a conditional demand model since nutritional status is modeled as a function of some inputs from the structural health production function and some variables from the reduced-form relations. Although this model cannot reveal all of the structural parameters nor the pathways through which particular policies may operate, it is informative about the effect of income and allows the direct effect on health of other variables, such as education, to be separated from any indirect effect that might operate through income.

7.1 Pooled OLS and probit models

The foundation of this analysis is a set of OLS and probit models that examine the relationship between nutritional status, as measured by height-for-age z-scores, and income and policy variables. Observations on all children who were under the age of 12 in each survey year are pooled and the panel is treated as if in each wave a new random sample was drawn from the population, thus exploiting only between-group variation. In the probit models, marginal effects are presented rather than probit coefficients[80]. Because preliminary models exhibit evidence of heteroskedasticity (according to the Breusch-Pagan test), because observations within community clusters are not independent and because (in panel data) there is likely to be serial correlation within communities and/or across individual observations over time, Huber community-clustered robust standard errors are used to avoid underestimating the true standard errors[81]. Since fertility may be endogenous (Schultz 1976; Wolfe and Behrman 1992), the real household income measure described in section 5 has been deflated by the total number of adults rather than total household size (Montgomery *et al.* 2000). The measure of income used is, therefore, the log of real per adult income.

Effect of income

The results, presented in Table 21, support the primary hypothesis of this paper. Real per adult income is found to be a significant correlate of improved child nutritional status in all models (see columns I and II), but the introduction of policy variables substantially mitigates its effect (see columns III-VI). The coefficient is largest in the basic income model (0.1 in the OLS model and -0.029 in the probit model), falls

[80] The marginal effects that are calculated are the average of the probabilities, rather than the probability of the average that is computed by STATA's "dprobit" or "mfxcompute" commands. There are a number of reasons why the probability of the averages method was avoided. First, the "probability of the average" method calculates the marginal effect for an observation with average characteristics which is hard to interpret since, for dummy variables, there may be no such person in the population who embodies such characteristics. Second, the values calculated are incorrect for variables that appear as interaction terms, squared terms or higher order terms. Third, and most importantly, it obscures the distribution of marginal effects: marginal effects are different for each observation, and some have large and some have small marginal effects. By contrast, the "average of the probabilities" method estimates the probability for each observation twice, changes the characteristics of interest, and recalculates any interaction or higher order terms. As it turns out, however, in this sample these two methods yield very similar marginal effects that are the same at least until (and including) the second decimal place.

[81] The heteroskedasticity-corrected standard errors are only slightly larger than the standard errors in the pooled model, but adjusting for clustering at the community-level enlarges standard errors substantially.

when being an only-child is controlled for (to 0.084 in the OLS models and -0.025 in the probit model), and falls still further when healthcare variables are introduced (to 0.064 and -0.019 respectively). The magnitude and direction of the coefficients of control variables are also in line with the hypotheses. The coefficients on the sex variable show that boys are not significantly more likely to be stunted than girls nor are their height-for-age z-scores significantly higher, implying that any differences in growth attainment are fully attributable to biological differences between the sexes and, thus, captured by the use of z-scores. The probability of stunting increases significantly as children age, but at a decreasing rate, indicating substantial age-related growth failure. Maternal education is significantly associated with favorable anthropometric outcomes, but only at the higher levels of educational attainment, specifically beyond primary school. This suggests that it is more than just basic information-processing skills or functional literacy – i.e. competencies associated with early education – that are critically important for child health in China. Also, since income is already controlled for in the model, the mechanism by which mother's education affects child nutritional status must also be independent of its effect via the income pathway[82]. Mother's height is also highly significant, with the smallest standard error of all variables, and the magnitude of its effect is very similar across all models.

Effect of the one child policy

When the one child policy variable is included (columns III and IV), it emerges that being an only-child is one of the most important predictors of nutritional status in the model. Indeed, only-children have height-for-age z-scores that are 0.43 deviations higher, on average, than children without siblings and probabilities of stunting that are 11 percentage points lower. As expected, including this variable exerts an attenuating effect on the magnitude of the effect of real per adult income. Its inclusion also weakens the coefficient on maternal education. This result is in line with expectations, first, because of the direct effect of education on the fertility decision – highly educated mothers tend to have fewer and healthier children (Schultz 1976)

[82] The role of father's education was also explored, but was not found to exert a significant independent effect on children's nutritional status. More importantly, adding father's education to the model had only a slight attenuating impact on the coefficients and marginal effects of mother's education. This means that the large coefficients on mother's education in the basic model were not capturing part of the effect of father's education as would be the case under an assortative matching hypothesis. Also, since controlling for father's education does not affect the marginal effect of income, it seems that the income effect in the simple model was not proxying for the unobserved cognitive competencies of the father.

– and, second, because of a possible spurious correlation between maternal education and number of children – the one child policy is more strictly enforced in urban areas (Scharping 2003) where educational opportunities are also better. The age effect is now also slightly reduced, as is expected, since, controlling for period effects, the probability of being an only-child is higher at younger age groups.

Effect of access to healthcare

In extending the model to measure the effect of access to healthcare (columns V and VI), variables capturing healthcare characteristics were constructed. In the CHNS, household heads listed all healthcare facilities that members of the household could use and provided information on the characteristics of each of those facilities. In order to avoid the endogeneity problem that the characteristics of healthcare facilities would determine which one is used by the household, this paper uses the characteristics of the healthcare facility that is closest to the household (in terms of traveling time) rather than the one that is most commonly used. Then, since the concern is with accessibility of healthcare in the community as a whole, rather than individual access to care, and to further reduce the potential correlation with unobserved household characteristics through simultaneity, five new variables were constructed that capture the non-self community mean values of each of the five facility characteristics, and their values were assigned to each household in the community. Missing values were imputed for those households with missing values on one or more healthcare characteristics (1,512 observations) by assigning the cluster/community mean value to these households. This reduced the total number of missing observations to 50[83]. Using the non-self cluster/community mean rather than the cluster/community mean carries the further advantage of introducing household-level variation into these variables since each observation within a cluster will take on a different value[84]. The five healthcare characteristics included are the non-self cluster means of one-way traveling time (in minutes), cost of travel (in real yuan)[85], waiting time (in minutes), typical cost of

[83] These 50 observations are all contained within the same community, indicating that healthcare information was missing for all households in the cluster/community, such that imputation of missing values is likely to have limited accuracy.

[84] As a consequence, potential collinearity in the community-level fixed effects model, estimated in section 7.2, is avoided.

[85] In order to reflect real prices, the price variables (i.e. the cost of treating a cold and cost of travel) are deflated using the same methodology used to deflate income.

treating a common cold at that facility (in real yuan) and medicine availability. Medicines are considered "available" if at least 90% of respondents in the community cluster report regular availability.

Of these healthcare variables, three emerge as significant at the 1% level, but the magnitudes of their effects are small. A higher time-cost of travel is associated with poorer nutritional status. However, contrary to expectations, higher financial costs of treatment and longer waiting times were not found to be *negatively* associated with better nutritional status, but rather *positively* and significantly associated with better nutritional status. A possible explanation for this result is that these two variables are, in fact, indicators of facility quality – people are willing to wait longer and pay higher prices for better quality healthcare which, in turn, is associated with improved nutritional status. Travel cost is not significant, but this is not surprising since only about 1% of the sample spends more than 1 real yuan (or 2 yuan nominal) to travel to health facilities; most people travel there on foot or by bicycle. The availability of medicine is not significant[86]. Together, all healthcare variables are jointly significant, and controlling for access to healthcare has the expected attenuating effect on income.

The positive coefficients on all three of the time dummies provide clear evidence of secular improvements in nutritional status in the first part of the decade, but with little change towards the end of the decade. For example, in the full model, controlling for changes in other time-varying factors, average height-for-age z-scores increased by more than one tenth of a standard deviation between 1991 and 1993 and by more than a quarter of a standard deviation between 1991 and 1997. Over the same two periods, the probability of stunting fell by 3 percentage points and 7 percentage points respectively. Thereafter, however, there appears to have been no further significant improvements in nutritional status: the null hypothesis that the 1997 and 2000 dummies are equal could not be rejected by the F-tests (OLS models) and chi-squared tests (probit models).

[86] The availability of medicine is not significant regardless of the cut-point used to define "availability": 50%, 60%, 70%, 80% or 90% of the time.

Table 21: Pooled OLS and pooled probit models of child nutritional status

	Income only		Income and one-child		Income, one-child and healthcare	
	I HAZ	II Stunting	III HAZ	IV Stunting	V HAZ	VI Stunting
Sex (Male=1)	-0.021	0.012	-0.031	0.015	-0.029	0.015
	-0.039	-0.015	-0.038	-0.014	-0.036	-0.014
Age (years)	-0.085	0.018	-0.055	0.01	-0.059	0.012
	(0.021)***	(0.008)**	(0.021)***	-0.007	(0.021)***	-0.007
Age squared	0.006	-0.001	0.004	-0.001	0.004	-0.001
	(0.001)***	(0.001)**	(0.001)***	(0.001)*	(0.001)***	(0.001)*
Maternal education (“no formal edication” is omitted)						
1-3 yrs primary	0.005	-0.024	-0.021	-0.019	-0.012	-0.016
	-0.102	-0.031	-0.1	-0.03	-0.095	-0.029
4-6 yrs primary	0.059	-0.033	0.038	-0.028	0.025	-0.02
	-0.089	-0.029	-0.087	-0.028	-0.082	-0.027
1-3 yrs lower middle	0.298	-0.098	0.245	-0.085	0.201	-0.068
	(0.081)***	(0.025)***	(0.078)***	(0.024)***	(0.074)***	(0.023)***
1-3 yrs upper or more	0.523	-0.161	0.416	-0.136	0.336	-0.106
	(0.098)***	(0.032)***	(0.091)***	(0.030)***	(0.085)***	(0.029)***
Mother’s height (cm)	0.063	-0.016	0.06	-0.015	0.056	-0.014
	(0.005)***	(0.002)***	(0.005)***	(0.002)***	(0.005)***	(0.002)***
Log real per adult income (yuan)	0.1	-0.029	0.084	-0.025	0.064	-0.019
	(0.026)***	(0.009)***	(0.024)***	(0.008)***	(0.024)***	(0.008)**
Year dummies						
1993	0.142	-0.031	0.143	-0.031	0.12	-0.022
	(0.026)***	(0.011)***	(0.025)***	(0.011)***	(0.033)***	(0.013)*
1997	0.279	-0.073	0.256	-0.066	0.271	-0.066
	(0.046)***	(0.017)***	(0.044)***	(0.017)***	(0.051)***	(0.019)***
2000	0.317	-0.064	0.24	-0.043	0.27	-0.047
	(0.063)***	(0.021)***	(0.060)***	(0.021)**	(0.073)***	(0.025)*
Index child is only-child			0.426	-0.113	0.325	-0.085
			(0.051)***	(0.017)***	(0.047)***	(0.017)***
Non-self cluster means of healthcare variables						
Cost of treating cold (real yuan)					0.015	-0.007
					(0.005)***	(0.003)**
Travel time (minutes)					-0.017	0.004
					(0.004)***	(0.001)***
Cost of travel (real yuan)					0.122	-0.041
					-0.107	-0.048
Waiting time (minutes)					0.007	-0.002
					(0.002)***	(0.001)***
Medicine regularly available					-0.032	0.003
					-0.052	-0.017
Constant	-11.816		-11.329		-10.513	
	(0.779)***		(0.739)***		(0.692)***	
Observations	6249	6249	6249	6249	6199	6199
R-squared	0.162		0.186		0.204	
Wald Test for equality of year dummies						

1993=1997						
F(1,175) =	9.39***		6.86***		10.88***	
Chi2(1) =		5.58***		3.96***		5.98***
1997=2000						
F(1,175) =	0.57		0.11		0	
Chi2(1) =		0.29		1.87		0.289
Wald test for joint significance of healthcare variables						
F(5, 175)					10.63***	
Chi2(5)						28.54***
Robust standard errors, clustered at the community level, in parentheses						
*significant at 10%; **significant at 5%; ***significant at 1%						

Source: China Health and Nutrition Survey, 1991-2000
Note: In the probit models, marginal effects are shown

7.2 Random effects and fixed effects results

A pooled OLS model is an inefficient estimator. Because the time-invariant error component v_j is in the composite error term in each time period, the composite error terms ε_{ijt} are serially correlated and, because the usual pooled OLS standard errors ignore this correlation, they will be incorrect.

If all of its assumptions hold, a random effects (RE) model would be the most efficient estimator. Indeed, the estimation of a random effects model with community-level random effects (see Column II) produces smaller coefficients than the OLS model on key parameters of interest, i.e. on the income and only-child variables (see Table 22, Columns I and II). The results of the Breusch-Pagan Lagrangian multiplier test [chi2(1)=872.06] indicate that these differences are significant, meaning that the unobserved effects are relatively important and the random effects model is preferred to the pooled OLS model.

However, the random effects model suffers from heteroskedasticity (according to the Breusch-Pagan test) and serial error correlation (according to the AR(1) test for autocorrelation). This means that the error correlations in the data do not appear to be of the same structure as is assumed by the random effects estimator, and the random effects estimates are likely to be inconsistent. Indeed, the key assumption of the random effects model, namely that the unobserved (time-invariant) community factors are uncorrelated with all the explanatory variables, seems theoretically implausible. More likely is that time-invariant factors are correlated with the explanatory variables – the assumption that underlies community fixed effects (FE) models. One such time-invariant factor could be the general level of economic development of the community which could exert a direct effect on nutritional status, but also be correlated with adult income, the number of children and the characteristics of healthcare facilities. In particular, the risk of bias posed by the potential endogeneity of the placement and characteristics of healthcare services is problematic. If the placement and/or characteristics of healthcare facilities are related to other unobserved factors that could also be correlated with nutritional status, such as population density, level of economic development and administrative rank, the coefficients of the healthcare variables may be biased[87]. To the extent that these unobserved factors are time-invariant, though, their influence can be negated when a community fixed effects model is used.

Specification tests support the use of the fixed effects model rather than the random effects estimator. The results of the Hausman test [chi2(18)=118.46] suggests that the random effects model, which is more efficient under the null hypothesis, will be biased if used with these data and should be rejected in favor of the more consistent community fixed effects model. Indeed, a comparison of the coefficients of the fixed effects model (shown in Table 22, column III) and the random effects model (Table 22, column II) reveals that the exclusion of community fixed effects did indeed bias the coefficients (in the predicted direction), implying that the explanatory variables are correlated with the error term[88]. Fixed effects are strongly and significantly positive [F(175, 6005)=5.07]. The rejection of the random effects model in favor of the fixed effects model implies that the community fixed effects model will also be preferred to the pooled OLS model since the random effects model is more efficient than the pooled OLS model.

The coefficients of the variables of interest are much smaller in the FE model than in the OLS and RE models. Both income and only-child coefficients are approximately half the size that they were in the RE model and one-third of the size that they were in the OLS model. Income is no longer significant. The real cost of treatment, travel time and waiting time have also lost their significance.

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[87] Gragnolati (1999), for example, enumerates three theories that could explain why the placement of health services is non-random: (i) theories of altruism, by which public resources would be allocated towards the more disadvantaged areas, (ii) pressure groups theories that predict allocation towards areas with higher demand and more lobbying power, and (iii) efficiency criteria and a concern with externalities that result in services being placed so as to maximize the overall well-being of the population.

[88] Since the Breusch-Pagan test shows evidence of heteroskedasticity in the fixed effects models [F(18,6180)=9.01], standard errors are adjusted for heteroskedasticity. Also, although in fixed effects models the time-invariant error component no longer causes the error terms to be clustered, errors may still be correlated if there are omitted community-specific time trends and, so, ex-post Huber standard errors are used.

7.3 Instrumental variables approach

A central econometric problem in the analysis is that the income variable is potentially endogenous due to the joint determination of income and child nutritional status through the labor/time allocation decision in the solution of the utility maximization problem. In other words, the labor/time allocation decision affects both income, via its effect on (especially female) labor supply, *and* child nutritional status, through influencing the time available for child care activities. For example, mothers with less healthy children may devote more time to caring for children at home as well as spend less time at work and on income-generation activities. Failure to control for this simultaneity will generate biased estimates of income, and also of other covariates that are correlated with income. Another potential source of endogeneity is measurement error in income. While simultaneity will tend to bias OLS estimates upwards, measurement error will tend to bias the estimates towards zero.

In an attempt to overcome potential endogeneity bias, income is now predicted using an instrumental variables (IV) approach to the OLS and fixed-effects models. Finding instruments that are simultaneously strong predictors of income in the first stage regression and also satisfy the exclusion criteria (i.e. are uncorrelated with the error term of the main equation) is challenging. There is also a trade-off between the consistency of IV estimates and the efficiency of OLS estimates: while an IV approach reduces the problems of bias introduced by endogeneity, it causes the variance of the parameters to increase. After giving consideration to a number of different instruments that have been used in previous studies of child nutritional status[89], two instruments – housing floorspace (measured in square meters) and the number of salary earners in the household – were selected.

In the instrumental variables OLS model (see Table 22, Column IV), these instruments were found to (i) be individually and jointly significant (F-test of joint significance equivalent to $F(2,156)=60.28$) in the first stage regression; (ii) together with other covariates, explain a sufficient share of the variation in income in the first stage regression (R-squared of 0.165); and (iii) be validly excluded from the main nutrition

[89] These include various combinations of the ownership or value of assets such as consumer durables, housing, equipment associated with household farms or enterprises, savings, land holdings and livestock ownership (see, for example, the wide variety used in the cross-country study of Haddad *et al.* 2003); community wages (see, for example, Attanasio *et al.* 2004) and, indicators of the industry or occupation of income earners (see, for example, Case *et al.* 2002).

equation in the instrumental variables model[90]. The Hausman test recommends rejecting the pooled OLS estimates in favor of the IV estimates, confirming that income was indeed endogenous in the OLS model. In the IV model, the income coefficient is significant and very large (0.46) – nearly seven times as large as the income coefficient in the original pooled OLS model – supporting the hypothesis that there was substantial underestimation of the income coefficient due to measurement error. The only-child variable, waiting times at health facilities and travel times to health facilities are significant and have magnitudes that are very similar to the original OLS model.

Once community-level fixed effects are controlled for[91], however, the significance of (instrumented) income disappears, as do the effects of the healthcare variables. The only-child variable remains significant – and large.

[90] All tests indicate that the exclusion criteria are satisfied. In the OLS model, the Hansen J test of overidentification could not be rejected (P>chi2=0.197). Regressions of the residuals from the main equation on the instruments and other covariates did not yield significant coefficients on the instruments, suggesting that the instruments are validly excluded from the main equation.

[91] In the fixed-effects model with instrumental variables, like the OLS model, the instruments (i.e. floorspace and the number of salary earners) are individually and jointly significant, explain a substantial share of the variation in income in the first stage regression and can be validly excluded from the main nutrition equation.

Table 22: Comparison of the results of the pooled OLS, RE, FE and instrumental variables models

	I	II	III	IV	V
	OLS	RE	FE	IV OLS	IV FE
Dependent variable: height for age z-scores					
Sex (Male=1)	-0.029	-0.012	-0.002	-0.02	0.001
	-0.036	-0.026	-0.033	-0.036	-0.026
Age (years)	-0.059	-0.068	-0.075	-0.083	-0.081
	(0.021)***	(0.019)***	(0.021)***	(0.021)***	(0.020)***
Age squared	0.004	0.005	0.005	0.005	0.005
	(0.001)***	(0.001)***	(0.001)***	(0.001)***	(0.001)***
Maternal education (“no formal education” is omitted)					
1-3 yrs primary	-0.012	0.017	0.041	-0.05	0.044
	-0.095	-0.062	-0.088	-0.098	-0.063
4-6 yrs primary	0.025	0.024	0.023	-0.033	0.021
	-0.082	-0.047	-0.076	-0.088	-0.048
1-3 yrs lower middle	0.201	0.152	0.118	0.081	0.105
	(0.074)***	(0.044)***	(0.066)*	-0.083	(0.047)**
1-3 yrs upper or more	0.336	0.251	0.173	0.188	0.159
	(0.085)***	(0.053)***	(0.081)**	(0.093)**	(0.057)***
Mother’s height (cm)	0.056	0.05	0.047	0.054	0.046
	(0.005)***	(0.003)***	(0.004)***	(0.004)***	(0.003)***
Log real per adult income (yuan)	0.064	0.045	0.021	0.462	0.138
	(0.024)***	(0.019)**	-0.022	(0.087)***	-0.087
Year dummies					
1993	0.12	0.127	0.148	0.137	0.149
	(0.033)***	(0.033)***	(0.027)***	(0.037)***	(0.034)***
1997	0.271	0.288	0.314	0.281	0.31
	(0.051)***	(0.039)***	(0.046)***	(0.051)***	(0.040)***
2000	0.27	0.296	0.34	0.377	0.357
	(0.073)***	(0.050)***	(0.063)***	(0.070)***	(0.054)***
Index child is only-child	0.325	0.208	0.104	0.311	0.119
	(0.047)***	(0.035)***	(0.046)**	(0.049)***	(0.039)***
Non-self cluster means of healthcare variables					
Cost of treating cold (real yuan)	0.015	0.014	0.007	0.005	0.003
	(0.005)***	(0.004)***	-0.005	-0.005	-0.003
Travel time (minutes)	-0.017	-0.005	0.003	-0.016	0.006
	(0.004)***	(0.003)**	-0.003	(0.004)***	-0.005
Cost of travel (real yuan)	0.122	0.166	0.158	0.133	0.161
	-0.107	(0.083)**	-0.122	-0.133	(0.088)*
Waiting time (minutes)	0.007	0.006	0.002	0.006	0.003
	(0.002)***	(0.001)***	-0.002	(0.002)***	-0.002
Medicine regularly available	-0.032	0.032	0.064	-0.018	0.055
	-0.052	-0.036	-0.042	-0.053	-0.039
Constant	-10.513	-9.533	-8.771	-12.895	-9.519
	(0.692)***	(0.436)***	(0.607)***	(0.796)***	(0.707)***
Observations	6199	6199	6199	6199	6199
R-squared	0.204		0.306		

Robust standard errors, clustered at the community level, in parentheses ; 176 community fixed effects
* significant at 10%; ** significant at 5%; *** significant at 1%

Source: China Health and Nutrition Survey, 1991-2000

7.4 Has the effect of income on nutritional status become more or less pronounced over time?

The models presented in the previous section assumed that the effect of income on child nutritional status has remained constant over time. However, in light of the extensive market liberalization of the 1980s and 1990s, it is possible that over the course of the decade income – i.e. the resources with which to purchase goods in the market – may have become an increasingly important determinant of nutritional status. In the healthcare market, for example, as was discussed in section 6.2, decreasing insurance coverage and increasing user fees suggests that the ability to pay fees may have increasingly determined access to services. Another example is the government food pricing policy which, since 1988, gradually abolished government grain procurement and urban rationing systems (Guo *et al.* 1999) so that the ability to pay increasingly became the only constraint on the purchase of food inputs.

To test this hypothesis, changes in the (partial) effect of income on child nutritional status (as measured by height-for-age z-scores) is examined by the introduction of a set of income-year interaction terms into the model[92]. The results, presented in Table 23 (columns I and II) reveal that the effect of income on nutritional status did indeed change over time, but not in the way that was hypothesized. The effect of income was, in fact, significantly *lower* in both 1993 and in 1997 than in 1991, regardless of whether time-invariant unobservables are controlled for (in the fixed effects models) or not (pooled OLS models). There is no significant difference between the 1991 and 2000 coefficients. Wald tests show that the income-time interaction terms for 1993 and 1997 are jointly significant (although not significantly different from each other), confirming that the income effect in 1993 and 1997 was significantly less pronounced than in 1991. Thus, on the whole, it seems that over the decade, income did not become a significantly more (or less) important determinant of nutritional status, but that there was a period in the first half of the decade when its importance declined.

[92] Ideally, the partial effects of the covariates on stunting would also be presented, but as Ai and Norton (2003) point out, in nonlinear models, the interaction effect, standard error and z-statistic change for each observation. Thus, the marginal effect of a change in both interacted variables is not equal to the marginal effect of changing just the interaction term; the sign may be different for different observations; and the statistical significance cannot be determined from the z-statistic reported in the regression output. Instead, to compute the magnitude of the interaction effect, one should compute the cross derivative of the expected value of the dependent variable, and the test for significance of the interaction effect must be based on the estimated cross partial derivative, not on the coefficient of the interaction term. Such results are impossible to present in a table.

Finally, it is worth pointing out that the observed effect of income is not robust to year-specific analysis, either in OLS models of height-for-age z-scores (see Table 23, columns III-VI) or in probit models of stunting (see Appendix, Table 26). This suggests that the effect of income on nutritional status observed in the pooled models (estimated in section 7.2) seems to have been driven by strong relationships of similar magnitude at the start and end of the decade (i.e. in 1991 and 2000 waves). Alternatively, if the downward bias on the income coefficients in the OLS and FE models in Table 22 was indeed largely attributable to measurement error, the failure to find a significant effect of income in 1993 and 1997 may reflect that there was more measurement error on the income variable in the 1993 and 1997 waves than in the 1991 and 2000 waves. Indeed, this hypothesis is supported by the results of a set of year-specific models in which real log per adult income is instrumented by housing floor space and number of salary earners: income coefficients that are significant at the 5% level are produced in 1991, 1993 and 1997 (although not in 2000) (see Appendix, Table 27). These year-specific models illustrate how different the conclusions about the impact of income on nutritional status would have been if only a single cross-section of data collected at one point during the 1990s had been examined. By contrast, coefficients on the only-child variables are strong and significant in all waves, and waiting times and travel times to healthcare facilities have a significant, if small, effect in almost all waves.

Table 23 Income-year interaction models and year-specific models

	Interactions		Year-specific models			
	OLS I	FE II	OLS 1991 III	OLS 1993 IV	OLS 1997 V	OLS2000 VI
Dependent variable: height for age z-scores						
Sex (Male=1)	-0.028	0	-0.052	0.015	-0.016	-0.071
	-0.036	-0.033	-0.046	-0.051	-0.06	-0.072
Age (years)	-0.059	-0.075	-0.09	-0.038	-0.079	0.085
	(0.021)***	(0.021)***	0.038**	-0.047	(0.047)*	-0.069
Age squared	0.004	0.005	0.007	0.002	0.005	-0.003
	(0.001)***	(0.001)***	0.003**	-0.003	-0.003	-0.005
Maternal education (“no formal education” is omitted):						
1-3 yrs primary	-0.012	0.04	0.02	0.016	-0.01	-0.3
	-0.096	-0.089	-0.11	-0.149	-0.161	-0.241
4-6 yrs primary	0.025	0.024	0.033	0.073	-0.004	-0.155
	-0.082	-0.076	-0.087	-0.113	-0.122	-0.231
1-3 yrs lower middle	0.207	0.123	0.181	0.242	0.155	0.142
	(0.073)***	(0.066)*	0.074**	(0.101)**	-0.107	-0.221
1-3 yrs upper or more	0.342	0.179	0.25	0.333	0.433	0.314
	(0.085)***	(0.081)**	0.093***	(0.116)***	(0.114)***	-0.238
Mother’s height (cm)	0.055	0.047	0.056	0.05	0.059	0.061
	(0.005)***	(0.004)***	0.006***	(0.008)***	(0.006)***	0.009***
Log real per adult income (yuan)	0.158	0.078	0.154	-0.002	-0.043	0.122
	(0.038)***	(0.037)**	0.040***	-0.044	-0.043	0.050**
Year dummies						
1993	1.288	0.782				
	(0.382)***	(0.372)**				
1997	1.634	1.462				
	(0.435)***	(0.397)***				
2000	0.35	0.124				
	-0.419	-0.395				
Income-year interaction terms						
Income*1993	-0.159	-0.086				
	(0.052)***	(0.050)*				
Income*1997	-0.184	-0.155				
	(0.059)***	(0.053)***				
Income*2000	-0.01	0.03				
	-0.058	-0.054				
Index child is only-child	0.325	0.106	0.247	0.405	0.353	0.207
	(0.047)***	(0.046)**	0.065***	(0.088)***	(0.082)***	0.080**
Non-self cluster means of healthcare variables						
Cost of treating cold (real yuan)	0.015	0.006	0.036	0.019	0.006	0.008
	(0.005)***	-0.005	0.014**	-0.014	-0.007	-0.006
Travel time (minutes)	-0.017	0.003	-0.019	-0.016	-0.024	-0.012
	(0.003)***	-0.003	0.004***	(0.006)***	(0.009)***	-0.01
Cost of travel (real yuan)	0.106	0.138	-0.012	0.337	0.107	-0.127
	-0.104	-0.118	-0.252	(0.148)**	-0.176	-0.148

Waiting time (minutes)	0.007	0.002	0.006	0.005	0.015	0.02
	(0.002)***	-0.002	0.002***	(0.003)*	(0.004)***	0.006***
Medicine regularly available	-0.023	0.074	-0.057	-0.004	-0.031	0.05
	-0.052	(0.042)*	-0.067	-0.103	-0.085	-0.102
Constant	-11.16	-9.202	-11.025	-9.067	-9.861	-12.023
	(0.710)***	(0.647)***	0.840***	(1.317)***	(0.990)***	1.553***
Observations	6199	6199	2173	1823	1373	830
R-squared	0.207	0.308	0.204	0.163	0.206	0.263
Wald tests for equality of parameters						
Income*1993=Income*1997 F(1,175)=	0.17	1.59				
Wald tests for joint significance of parameters						
Income*1993 and Income*1997 F(2,175) =	7.02***	4.39***				
Income*1997 and Income*2000 F(2,175) =	6.70***	7.90***				
Income*1993 and Income*1997 and Income*2000 F(3,175) =	6.52***	5.88***				

Robust standard errors, clustered at the community level, in parentheses
* significant at 10%; ** significant at 5%; *** significant at 1%

Source: China Health and Nutrition Survey, 1991-2000

Note: Results for year-specific stunting models are presented in the Appendix, Table 8.

7.5. Do the effects of policy variables vary by income and gender?

It may be that the nutritional impact of being an only-child and having access to healthcare is more salient for children with certain characteristics, and in particular for children that belong to relatively disadvantaged subgroups. To examine this, interaction terms are introduced that capture how the effect of being an only-child and having access to healthcare varies by income and gender.

With respect to income, neither policy variable is regressive or progressive (see Appendix, Tables 28 and 30). Being an only-child does not have a statistically different effect on the nutritional status of children in households with higher incomes than on the nutritional status of children in households with lower incomes. The interaction of income and healthcare also produces insignificant coefficients. This implies that the extent to which the availability, quality and cost of healthcare will affect child nutritional status does not depend on the amount of income that the household has available with which to purchase that care.

The policies appear to be largely gender-neutral, too (see Appendix, Tables 28 and 29). While being an only child was a highly significant variable in all models estimated in this paper, boys who are only-children do not have z-scores that are significantly different from girls who are only-children. Distance to the healthcare facility (as measured by traveling time) is the only healthcare variable for which a statistically significant gender effect is observed. Since longer traveling times are associated with poorer nutritional status (i.e. lower height-for-age z-scores), the positive coefficient on the interaction term shows that the association is somewhat weaker for boys than for girls. This suggests that distance to the healthcare facility may be less of an obstacle to seeking care when the child is a boy than a girl, and perhaps indicative of a household preference for investment in son's health over daughter's health.

8. Conclusion

This paper sought to explain the variation in child nutritional status in China during the 1990s. Following the predictions of a theoretical household utility-maximizing model, it was hypothesized that the variation in child nutritional status might be attributable to variation in household income. Indeed, OLS and probit models with data pooled over four time periods produced income coefficients that were large and significant.

Since the theoretical pathway from income to nutritional status is influenced by policy interventions, the paper also explored the extent to which the effect of income was mediated by the outcomes of two policies: first, the one child policy and its effect on whether or not the index child is an only-child, and, second, the healthcare transition and the changes it brought about in the characteristics of the healthcare facilities to which households in a community have access. Not only were these policy variables found to be significant, but when they were incorporated into the model, the effect of income fell sharply. Income still remained a significant correlate of child nutritional status, though.

Once unobserved time-invariant community characteristics were controlled for in a community fixed effects model, the effect of income disappeared entirely. The same result was obtained using an instrumental variables approach: the income effect was large and significant in initial pooled OLS and probit models – even when policy variables were included – but when community fixed-effects were introduced, the income coefficients became insignificant.

These results should not be interpreted as evidence that income does not affect on nutritional status nor that policies aimed at improving household incomes may not be effective in improving child nutritional status. Since the introduction of community fixed effects-absorbs inter-community variation, it is the intra-community variation in income that is being considered in a fixed-effects estimation. If there is not sufficient variation on the income variable across time *within* communities, there will not be a significant measured effect of income, The fact that the income effect disappears when time-invariant community fixed-effects are controlled for does not necessarily imply that income is not an important determinant of nutritional status, but may simply imply that most of the relationship between income and nutritional status observed in the initial models is attributable to variation in income across communities rather than between households within communities.

Access to healthcare, measured by five characteristics of healthcare facilities, was found to be an important correlate of child nutritional status, at least in OLS and probit models. Specifically, an increase in the time taken to travel to healthcare facilities was associated with worse nutritional outcomes and, if the cost of treatment and waiting time can be assumed to be indicators of facility quality, then better facility quality was associated with better nutritional status. The financial cost of care was not significant, however. This suggests that the concern that the increasing user fees and decreasing insurance coverage associated with the healthcare transition may have had detrimental consequences for child nutritional status, by increasing the cost of treatment, may not have as strong a foundation as was hypothesized. The location of facilities, specifically how long it takes to travel to a healthcare facility, and the quality of care do appear to be important factors, though. However, one should be careful not to overstate the importance of healthcare characteristics: once time-invariant community factors are controlled for in a fixed effects model, the effects of the healthcare variables are no longer significant. This implies that in the OLS model, these health variables may have been picking up other types of variation across communities that are correlated with variation in community health facility characteristics. As such, they may simply reflect other community-level changes, such as generalized improvements in transportation and infrastructure.

Further, while the results of this analysis should not be interpreted as a measure of the effect of the one child policy *per se* (since the variable used in the analysis only captures one outcome of the policy), the analysis is illuminating with respect to the nutritional consequences of reduced fertility. Controlling for the socioeconomic characteristics of the household, only-children had height-for-age z-scores that were a third (OLS models) to a fifth (FE models) of a standard deviation larger than that of children with siblings. While the results are consistent with a conclusion that any policy measure that can effectively encourage parents to have only one child could have large positive consequences for child nutritional status, it is likely that the difference in nutritional status between only-children and children with siblings in China is larger than one can expect in other countries where reduced fertility is the result of other policies, such as the promotion of family-planning methods and the enhancement of female autonomy. This is because the particular rewards and sanctions associated with the one child policy have direct consequences for child health. For example, only-children often receive health subsidies and other health-promoting inputs that are denied to subsequent children, and women may fail to seek healthcare for unapproved higher order births out of fear of financial penalties.

Appendix: Additional figures and tables for Paper III

Derivation of the reduced form child health demand model

A one-period unitary model of decision-making is assumed whereby a household maximizes a **utility function** that depends on the consumption of commodities including a composite (market-purchased) household consumption good (X_{jt}), the consumption of leisure by all household members (L_{jt}) and child health or nutritional status (H_{ijt}), conditional on a set of household taste and preferences shifters (W_{jt}):

$$U_{jt} = U(X_{jt}, L_{jt}, H_{ijt}; W_{jt}) \qquad (1)$$

The household faces three constraints:

(i) a health production function representing the technology available to the household to transform available inputs into the health of the child. The health or nutritional status of child *i* in household *j* in period *t* can be described by the function

$$\mathrm{H}_{\mathrm{ijt}} = h(M_{ijt}, Z_{ijt}, V_{ijt}; G_{ijt}, F_{jt}) \qquad (2)$$

where nutritional status depends on M_{ijt}, the endogenous material health inputs (such as immunizations, nutritional intake, healthcare utilization); on Z_{ijt}, the exogenous characteristics of the child and the household in which it lives; and on V_{ijt}, the endogenous child health-related time inputs. Child health is also conditional on two vectors of exogenous child-specific (G_{ijt}) and household/community-specific variables (F_{jt}) that are time-varying or time-invariant indicators of health technology[93]. As such, they influence the choice of health inputs and/or the efficiency with which existing inputs are combined to produce child health, and in so doing affect the shape of the health production function. Exogenous child-specific factors (G_{ijt}) could include age and gender. An example of an exogenous endowment at the household-level (F_{jt}) is the education of the caregiver (typically the mother) who makes health-related decisions that impact on the child's health. At the community-level, prevailing cultural norms regarding child health practices, the

[93] F_{ijt} and G_{ijt} variables can be thought of as exogenous productivity shifters, which may be either permanent or temporary shocks (Currie 2000).

availability of healthcare facilities, relative food prices and communication or transportation infrastructure are examples of F_{jt} variables.

(ii) a **time endowment** reflecting the total available time available to the household (N_{jt}) to allocate between wage labor (T_{jt}), leisure (L_{jt}) and health-related activities (V_{jt})

$$N_{jt} = T_{jt} + L_{jt} + V_{jt} \tag{3}$$

(iii) the **budget constraint** giving the total financial resources, consisting of wage income ($w_t T_{jt}$), non-wage income (Y_{jt}) and net assets (A_{jt}), available to the household with which to purchase market-produced health-related inputs (M_{ijt}) and other consumption goods (X_{jt}) at prices p^M and p^X respectively

$$p_t^X X_{jt} + p_t^M M_{ijt} = w_t T_{jt} + Y_{jt} + A_{jt} \tag{4}$$

The budget constraint (4) is combined with the time endowment (3) to yield the **full income constraint**:

$$p_t^X X_{jt} + p_t^M M_{ijt} = w_t (N_{jt} - L_{jt} - V_{ijt}) + Y_{jt} + A_{jt} \tag{5}$$

Wages are assumed to be exogenous and can, thus, represent the predetermined opportunity cost of time.

Re-arranging the terms so that total consumption, of both goods and time, is subject to the budget constraint including time endowment gives:

$$p_t^X X_{jt} + p_t^M M_{ijt} + w_t L_{jt} + w_t V_{ijt} = w_t N_{jt} + Y_{jt} + A_{jt} \tag{6}$$

The left-hand side of (6) represents resource costs or expenditure on goods and services including leisure, which cannot exceed the total resources available to the household. Under the assumption that the utility function is increasing and quasi-concave and the underlying functions have desirable properties so that an internal maximum can be obtained, the utility function can be maximized subject to the technology and full income constraint to solve for the optimal amount of material health inputs (M_{ijt}*) and health-related time inputs (V_{ijt}*), as well as the optimal amount of the household consumption good (X_{jt}*) and leisure time (L_{jt}*) in each time period to obtain the following set of reduced-form demand functions.

$$M_{ijt}{}^* = m(p_t^X, p_t^M, w_t, N_{jt}, Y_{jt}, A_{jt}; G_{ijt}, F_{jt}) \tag{7}$$

$$V_{ijt}{}^* = v(p_t^X, p_t^M, w_t, N_{jt}, Y_{jt}, A_{jt}; G_{ijt}, F_{jt}) \tag{8}$$

$$X_{jt}* = x(p_t^X, p_t^M, w_t, N_{jt}, Y_{jt}, A_{jt}; G_{ijt}, F_{jt}) \quad (9)$$

$$L_{jt}* = l(p_t^X, p_t^M, w_t, N_{jt}, Y_{jt}, A_{jt}; G_{ijt}, F_{jt}) \quad (10)$$

The left hand side variables are all endogenous and the right hand side variables are all exogenous variables, i.e. prices, endowments and predetermined wealth.

The health demand function can be derived by substituting equations (7) and (8) into (2):

$$\mathrm{H}_{\mathrm{ijt}} = h(m(p_t^X, p_t^M, w_t, N_{jt}, Y_{jt}, A_{jt}; G_{ijt}, F_{jt}), Z_{ijt}, v(p_t^X, p_t^M, w_t, N_{jt}, Y_{jt}, A_{jt}; G_{ijt}, F_{jt}); G_{ijt}, F_{jt}) \quad (11)$$

The particular functional form *h*( ) depends on the underlying functions characterizing household preferences and the health production function.

Construction of the income measure

The income measure aggregates multiple cash income streams, including (i) wages, salaries and bonuses from wage-employment (estimated for an annual period and collected at the job level), (ii) cash income from various types of self-employment, including home gardening, home farms, collective farming activities, animal husbandry, household fishing, and small household business (based on monthly recall and collected at the animal, crop or business level), (iii) income flows from social security, rental of housing and durables, and (iv) remittances from friends and relatives. In this sample, non-wage income accounts for a substantial share of total income: in rural areas, income from the self-employment activities accounts for more than half of all income, while in urban areas, social security and rental incomes together account for almost a quarter of all income.

The market value of in-kind income, as estimated by the household head, is also included in the measure since failure to include flows of goods (in addition to flows of money) will underestimate the resources that the household has at its disposal to invest in the production of child health. For example, over and above the salary, many Chinese households receive in-kind benefits, such as food and gifts, from work units and local enterprises, or from other households.

Also, since some households engaged in agricultural, animal husbandry or fishing activities satisfy a large share of their food needs through consuming some or all of their crops, livestock or fish, the household head's estimate of the price he or she would have received for those items at market is included in the income measure.

The variable is constructed so as to reflect net values; all expenses associated with a particular income source are subtracted from the total revenue associated with that income source.

The income measure is also adjusted for temporal and spatial variation in prices. This accounts for the fact that households in a pooled sample that report the same income may face different prices in different periods, and in different geographic areas, so that the number of items (and thus utility) that can be purchased for a given income varies.

There is no published consumer price index for China (at least prior to 2000) with a clear base year. Instead, the China State Statistical Bureau's (SSB) publishes an annual consumer price index ratio that shows the shift in the cost of living within urban and rural areas in each province. Consequently, a special consumer price index has been developed specifically for use with the CHNS data. The first step in the construction of this index is to use a 57 item consumer basket[94], obtained from a 1989 income and expense survey in Chinese urban households (Ren *et al.* 1989), and urban prices from SSB data to create urban consumer costs for each province in the CHNS for the years 1989, 1991, 1993 and 1997. Next, an urban-rural ratio is calculated for each survey year, using food price survey averages (for the goods that appear in both the SSB and CHNS baskets) for rural and urban areas. Finally, using the cost of the consumer basket based on the SSB basket, price data and urban-rural ratio, the rural cost for the same consumer basket for each year can be calculated for each province, with urban Liaoning province for 1988 as the base year[95]. One advantage of this measure is that it captures prices variation at a very low level, i.e. that of the community. Another is that because prices are obtained from a community survey, the item chosen, and its quality, is not endogenous.

Table 24: Change in the mean of real *per capita* income across households in the sample, 1991-2000, in 1988 yuan

	1991	1993	1997	2000	% change 1991-2000
Total sample	1645	1290	1537	1401	-14.8
Urban	2036	1851	2014	1846	-9.3
Rural	1479	1074	1331	1211	-18.1
Jiangsu	1875	1727	2062	1996	6.5
Shandong	1631	1176	1587	1499	-8.1
Henan	1386	855	1272	1110	-19.9
Hubei	1928	1258	1372	1291	-33.0
Hunan	2044	1569	1987	1557	-23.8
Guangxi	1345	1372	1493	1361	1.2
Guizhou	1428	1157	1141	1108	-22.4

Source: China Health and Nutrition Survey, 1991-2000

[94] Included in this basket are various types of food products, fuels, cigarettes, soaps, fabric, clothing, and some consumer durables.

[95] Income data collected in 1989 reflect income in calendar year 1988 and, therefore, the deflator uses 1988 prices. Liaoning province is not one of the provinces in the analysis in this paper.

Table 25: Change in the percentage of only-children, under 12, 1991-2000

	1991	**1993**	**1997**	**2000**	**% change 1991-2000**
Total sample	27.1%	27.4%	33.6%	42.8%	57.9%
Urban	42.9%	40.8%	46%	55.1%	28.4%
Rural	21.3%	23%	29%	38.7%	81.7%
Male	27.9%	28.8%	34.1%	42.5%	52.3%
Female	26.2%	25.7%	33%	43.1%	64.5^
Jiangsu	51.2%	52.6%	60.2%	59.8%	16.8%
Shandong	40.7%	36.5%	41.5%	63.9%	57%
Henan	28.9%	29.8%	31.9%	38.7%	33.9%
Hubei	23.9%	21.9%	27.7%	40%	67.4%
Hunan	23.1%	23%	34.9%	48.8%	111.3%
Guangxi	18.6%	20.9%	24.4%	32.7%	75.8%
Guizhou	16.2%	19.4%	25.4%	34.1%	110.5%

Source: China Health and Nutrition Survey, 1991-2000

Table 26: Probit estimates of year-specific stunting models

	1991	1993	1997	2000
Dependent variable: Stunting=1				
Sex (Male=1)	0.029	0.007	0.017	-0.013
	-0.018	-0.02	-0.021	-0.027
Age (years)	0.031	0.011	0.01	-0.038
	(0.013)**	-0.016	-0.016	(0.020)*
Age squared	-0.003	-0.001	-0.001	0.002
	(0.001)***	-0.001	-0.001	-0.001
Maternal education ("no formal education" is omitted)				
1-3 yrs primary	-0.037	0.01	-0.02	-0.014
	-0.04	-0.046	-0.044	-0.067
4-6 yrs primary	-0.001	-0.049	-0.009	-0.025
	-0.031	-0.038	-0.04	-0.058
1-3 yrs lower middle	-0.063	-0.072	-0.048	-0.11
	(0.024)***	(0.033)**	-0.035	(0.055)**
1-3 yrs upper middle or more	-0.126	-0.079	-0.106	-0.124
	(0.034)***	(0.041)*	(0.046)**	(0.061)**
Mother's height (cm)	-0.017	-0.013	-0.012	-0.014
	(0.002)***	(0.003)***	(0.002)***	(0.003)***
Log real per adult income (yuan)	-0.051	0	0.018	-0.031
	(0.014)***	-0.015	-0.015	(0.014)**
Index child is only-child				
	-0.083	-0.117	-0.062	-0.042
	(0.025)***	(0.028)***	(0.026)**	-0.027
Non-self cluster means of healthcare variables				
Cost of treating cold (real yuan)	-0.017	-0.007	-0.007	-0.003
	(0.006)***	-0.006	(0.004)*	-0.003
Travel time (minutes)	0.005	0.004	0.008	0.002
	(0.001)***	(0.002)*	(0.003)**	-0.003
Cost of travel (real yuan)	0.038	-0.205	-0.022	0.073
	-0.049	(0.082)**	-0.056	-0.06
Waiting time (minutes)	-0.003	-0.001	-0.005	-0.005
	(0.001)***	-0.001	(0.002)***	(0.003)*
Medicine available	-0.003	-0.005	0.013	-0.024
	-0.023	-0.037	-0.031	-0.03
Observations	2173	1823	1373	830
Robust standard errors, clustered at the community level, in parentheses				
* significant at 10%; ** significant at 5%; *** significant at 1%				

Source: China Health and Nutrition Survey, 1991-2000

Note: Marginal effects are shown

Table 27: Year-specific OLS models with instrumental variables

	1991	1993	1997	2000
Dependent variable: height for age z-scores				
Sex (Male=1)	-0.033	0.037	-0.028	-0.072
	-0.043	-0.052	-0.06	-0.072
Age (years)	-0.101	-0.089	-0.098	0.082
	(0.038)***	(0.048)*	(0.050)*	-0.07
Age squared	0.007	0.005	0.006	-0.003
	(0.003)**	-0.003	(0.003)*	-0.005
Maternal education (“no formal education” is omitted)				
1-3 yrs primary	-0.001	-0.098	-0.027	-0.298
	-0.124	-0.15	-0.163	-0.238
4-6 yrs primary	-0.023	-0.025	-0.055	-0.155
	-0.096	-0.126	-0.137	-0.228
1-3 yrs lower middle	0.111	0.05	-0.002	0.136
	-0.084	-0.125	-0.137	-0.22
1-3 yrs upper middle or more	0.165	0.099	0.255	0.305
	-0.11	-0.136	(0.144)*	-0.237
Mother’s height (cm)	0.051	0.052	0.056	0.061
	(0.006)***	(0.008)***	(0.007)***	(0.009)***
Log real per adult income (yuan)	0.548	0.51	0.366	0.15
(instrumented)	(0.149)***	(0.148)***	(0.179)**	-0.128
	0.223	0.395	0.324	0.211
Index child is only-child	0.223	0.395	0.324	0.211
	(0.068)***	(0.091)***	(0.085)***	(0.082)**
Non-self cluster means of healthcare variables				
Cost of treating cold (real yuan)	-0.018	-0.014	-0.022	-0.012
	(0.004)***	(0.007)**	(0.010)**	-0.009
Travel time (minutes)	0.017	0	0.001	0.007
	-0.014	-0.015	-0.008	-0.007
Cost of travel (real yuan)	-0.061	0.177	0.257	-0.135
	-0.26	-0.199	-0.221	-0.148
Waiting time (minutes)	0.007	0.003	0.012	0.02
	(0.002)***	-0.004	(0.004)***	(0.006)***
Medicine available	0.007	0.012	-0.072	0.048
	-0.078	-0.118	-0.096	-0.102
Constant	-13.043	-12.738	-12.091	-12.189
	(1.071)***	(1.718)***	(1.234)***	(1.501)***
Observations	2173	1823	1373	830
R-squared	0.153	0.047	0.141	0.263

Robust standard errors, clustered at the community level, in parentheses
* significant at 10%; ** significant at 5%; *** significant at 1%

Source: China Health and Nutrition Survey, 1991-2000

Table 28: Regressions of height-for-age z-scores on only-child interaction terms

	Sex		Income	
	OLS	**FE**	**OLS**	**FE**
Sex (Male=1)	-0.024	-0.002	-0.029	-0.002
	-0.042	-0.037	-0.036	-0.033
Log real per adult income (yuan)	0.064	0.021	0.052	0.024
	(0.024)***	-0.022	(0.027)*	-0.024
Index child is only-child	0.336	0.102	-0.048	0.206
	(0.061)***	(0.061)*	-0.367	-0.348
Interaction terms				
Sex*Only-child	-0.019	0.003		
	-0.073	-0.072		
Income*Only-child			0.051	-0.014
			-0.05	-0.048
Constant	-10.521	-8.77	-10.412	-8.795
	(0.694)***	(0.607)***	(0.698)***	(0.608)***
Observations	6199	6199	6199	6199
R-squared	0.204	0.306	0.204	0.306
No. of community fixed effects		176		176
Robust standard errors, clustered at the community level, in parentheses				
*significant at 10%; **significant at 5%; ***significant at 1%				

Source: China Health and Nutrition Survey, 1991-2000

Note: Only select coefficients shown; other coefficients are the same as in the full pooled OLS and FE models (see, for example, columns I and III, Table 22)

Table 29: Regressions of height-for-age z-scores on interaction between gender and healthcare

	Treatment cost		Travel time		Travel cost		Waiting time		Medicine	
	OLS	**FE**	**OLS**	**FE**	**OLS**	**FE**	**OLS**	**FE**	**OLS**	**FE**
Sex (Male=1)	-0.051	-0.018	-0.115	-0.09	-0.029	-0.009	-0.035	-0.016	0.046	0.054
	-0.048	-0.044	(0.057)**	(0.054)*	-0.037	-0.034	-0.051	-0.047	-0.059	-0.059
Cost of treating cold (real yuan)	0.013	0.005	0.015	0.007	0.015	0.007	0.015	0.007	0.015	0.007
	(0.006)**	-0.005	(0.005)***	-0.005	(0.005)***	-0.005	(0.005)***	-0.005	(0.005)***	-0.005
Travel time (minutes)	-0.017	0.003	-0.021	-0.001	-0.017	0.003	-0.017	0.003	-0.017	0.003
	(0.004)***	-0.003	(0.004)***	-0.004	(0.004)***	-0.003	(0.004)***	-0.003	(0.004)***	-0.003
Cost of travel (real yuan)	0.121	0.158	0.116	0.154	0.123	0.097	0.122	0.159	0.122	0.158
	-0.107	-0.122	-0.104	-0.116	-0.119	-0.134	-0.107	-0.122	-0.107	-0.122
Waiting time (minutes)	0.007	0.002	0.007	0.002	0.007	0.002	0.007	0.002	0.007	0.002
	(0.002)***	-0.002	(0.002)***	-0.002	(0.002)***	-0.002	(0.002)***	-0.002	(0.002)***	-0.002
Medicine regularly available	-0.031	0.064	-0.032	0.063	-0.032	0.064	-0.032	0.064	0.021	0.102
	-0.052	-0.043	-0.052	-0.042	-0.052	-0.042	-0.052	-0.042	-0.059	(0.054)*
Sex*Cost	0.004	0.003								
	-0.006	-0.005								
Sex*Travel time			0.008	0.009						
			(0.004)**	(0.003)**						
Sex*Travel cost					-0.002	0.129				
					-0.119	-0.119				
Sex*Waiting time							0	0.001		
							-0.002	-0.002		
Sex*Medicine									-0.098	-0.072
									-0.06	-0.06
Observations	6199	6199	6199	6199	6199	6199	6199	6199	6199	6199
R-squared	0.204	0.306	0.204	0.307	0.204	0.306	0.204	0.306	0.204	0.306

Robust standard errors, clustered at the community level, in parentheses
*significant at 10%; **significant at 5%; ***significant at 1%

Source: China Health and Nutrition Survey, 1991-2000

Note: Only select coefficients shown; other coefficients are the same as in the full pooled OLS and FE models (see, for example, columns I and III, Table 22)

Table 30: Regressions of height-for-age z-scores on interaction between income and healthcare

	Treatment cost		Travel time		Travel cost		Waiting time		Medicine	
	OLS	**FE**	**OLS**	**FE**	**OLS**	**FE**	**OLS**	**FE**	**OLS**	**FE**
Log real per adult income (yuan)	0.102	0.025	0.075	-0.007	0.07	0.028	0.095	0.031	0.047	0.007
	(0.034)***	-0.028	(0.040)*	-0.034	(0.026)***	-0.023	(0.032)***	-0.026	-0.044	-0.042
Cost of treating cold (real yuan)	0.073	0.012	0.015	0.007	0.016	0.007	0.016	0.007	0.015	0.007
	(0.037)*	-0.031	(0.005)***	-0.005	(0.005)***	-0.005	(0.005)***	-0.005	(0.005)***	-0.005
Travel time (minutes)	-0.017	0.003	-0.009	-0.016	-0.017	0.003	-0.017	0.003	-0.017	0.003
	(0.004)***	-0.003	-0.022	-0.02	(0.004)***	-0.003	(0.004)***	-0.003	(0.004)***	-0.003
Cost of travel (real yuan)	0.142	0.159	0.121	0.165	0.705	0.723	0.126	0.159	0.121	0.158
	-0.107	-0.122	-0.106	-0.126	-0.519	-0.612	-0.106	-0.122	-0.107	-0.122
Waiting time (minutes)	0.007	0.002	0.007	0.002	0.007	0.002	0.025	0.007	0.007	0.002
	(0.002)***	-0.002	(0.002)***	-0.002	(0.002)***	-0.002	(0.013)*	-0.012	(0.002)***	-0.002
Medicine regularly available	-0.032	0.064	-0.033	0.067	-0.032	0.064	-0.033	0.064	-0.197	-0.082
	-0.051	-0.042	-0.052	-0.042	-0.052	-0.042	-0.052	-0.042	-0.388	-0.347
Income*Cost	-0.008	-0.001								
	(0.005)*	-0.004								
Income*Travel time			-0.001	0.003						
			-0.003	-0.003						
Income*Travel cost					-0.079	-0.077				
					-0.069	-0.075				
Income*Waiting time							-0.002	-0.001		
							-0.002	-0.002		
Income*Medicine									0.022	0.02
									-0.052	-0.046
Observations	6199	6199	6199	6199	6199	6199	6199	6199	6199	6199
R-squared	0.204	0.306	0.204	0.306	0.204	0.306	0.204	0.306	0.204	0.306

Robust standard errors, clustered at the community level, in parentheses
*significant at 10%; **significant at 5%; ***significant at 1%

Source: China Health and Nutrition Survey, 1991-2000

Note: Only select coefficients shown; other coefficients are the same as in the full pooled OLS and FE models (see, for example, columns I and III, Table 22)

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