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Young Children Left Behind in Rural China

Alan de Brauw

Ren Mu

Markets, Trade and Institutions Division

INTERNATIONAL FOOD POLICY RESEARCH INSTITUTE

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AUTHORS

Alan de Brauw, International Food Policy Research Institute
Senior Research Fellow, Markets, Trade and Institutions Division
a.debrauw@cgiar.org

Ren Mu, Texas A&M University
The Bush School of Government and Public Service
rmu@tamu.edu

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ABSTRACT

The unprecedented, large-scale, rural-to-urban migration in China has left many rural children living apart from their parents. Yet the consequences for child development of living without one or more parents due to migration are largely unknown. In this study, we examine the impact of parental migration on one measure of child development, the nutritional status of young children in rural areas. We use the interaction terms of wage growth in provincial capital cities with initial village migrant networks as instrumental variables to account for migration selection. Our results show that parental migration has no significant impact on the height of children but that it improves their weight. We provide suggestive evidence that the improvement in weight may be achieved through increased access to tap water in households with migrants. To conclude the paper, we raise concerns about the sustainability of the impact.

Keywords: migration, children, nutrition, rural China

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1. INTRODUCTION

The migration of labor out of agriculture is a primary feature of the economic development process, and the study of migration as an economic process has a long history (Lewis 1954; Fei and Ranis 1964). The role of migration in the economy has recently regained prominence in policy discussions and in the research community because of an increased flow of labor both internationally and within large countries, such as India and China. However, Clemens (2011) notes that much research has focused on the effects of immigration while neglecting emigration, the effects of migration on those left behind.

From a microeconomic perspective, the effects of emigration on source households and communities can be complex. Migrants typically continue to have economic interactions with the households and communities they leave behind (Stark and Bloom 1985), and these interactions can be particularly important when markets do not function well. For example, migrant remittances can help households overcome credit constraints that hinder investment in the human capital of the next generation (Beine, Docquier, and Rapoport 2008; Yang 2008). When migrants leave households, however, the time investment in raising children also decreases (de Brauw and Mu 2011); moreover, as household members live apart because of migration, they cannot perfectly monitor one another's behavior. Consequently, the adults remaining in the village may not serve the best interests of the children living with them (Chen 2006). Consequently, migration can affect human capital outcomes of children in several ways, and the overall effect depends on the relative importance of these factors.

Focus on children's nutrition is important because poor nutrition outcomes may have long-term consequences for children. Multiple studies have found children's health and nutritional status to have a sizable and statistically significant positive impact on educational outcomes in many developing countries, such as Pakistan (Alderman et al. 2001), Ghana (Glewwe and Jacoby 1995), Guatemala (Maluccio et al. 2009), and Kenya (Miguel and Kremer 2004).[1] Moreover, positive impacts seem to occur among both boys and girls (Glewwe and Miguel 2008).

Several recent studies suggest the potential importance of migration for child nutrition among those left behind. Gibson, McKenzie, and Stillman (2011) find that in Tonga, height-for-age *z* scores (HAZ) are lower among children younger than 18 years old left behind by migrants to New Zealand.[2] They find no impact on weight-for-age *z* scores (WAZ) or BMI-for-age *z* scores (BMIZ).[3] Carletto et al. (2011) find a positive impact on HAZ scores among Guatemalan children left behind by emigrants to the United States. Evidence from Tajikistan also shows that children in communities with more migrants have higher HAZ-scores (Azzarri and Zezza 2011). Mansuri (2006) and Nobles (2007) study the effects of international migration from Pakistan and Mexico, respectively, on child HAZ-scores, with the former finding a positive effect and the latter a negative one. Differences in data, country contexts, child age definitions, empirical specifications, and methods all may have contributed to the differences in these results, but they also highlight the complex relationship between migration and child nutrition, a subject that deserves further investigation.

China is a particularly salient place to study this issue for several reasons. First, the flow of internal migrant labor has increased rapidly since economic reforms began and in particular since labor market restrictions were loosened (Liang and Ma 2004; Fan 2008), yet partially as the result of the household registration (*hukou*) system in China, migrants are still largely barred, either legally or financially, from accessing education for their children in urban areas. Poor housing conditions for migrant workers in urban areas also prevent family migration (World Bank 2009); consequently, whole-family migration is rare, and many young and school-age children stay behind in rural villages. Recent

[1] For a literature review of recent studies on the impact of child health and nutrition on education in developing countries, please see Glewwe and Miguel (2008).

[2] They also find that children who accompany migrants are healthier using several other measures of health status.

[3] BMI stands for body mass index, calculated by dividing a person's body weight in kilograms by his or her height in meters squared (weight [kg]/height$[m]^2$). BMI-for-age *z* scores are used like weight-for-height *z* scores but also control explicitly for age.

estimates suggest that as many as 23 million children (younger than age 14) are left behind in migrant-sending regions while their parents work elsewhere (Guo 2009).

This study builds on our previous analysis of the relationship between migration and nutritional status among children in rural China (de Brauw and Mu 2011). We previously found that children aged 7 to 12 years in migrant households are more likely to be underweight perhaps because they spend more time doing household chores than children in nonimmigrant households. We also found that younger children (aged 2 to 6 years) in migrant households are less likely to be overweight when no grandparent lives with them. In this study, we examine distribution means of HAZ, BMIZ, and WAZ scores to complement the previous findings on measures based on the tails of the BMIZ distribution—overweight and underweight status. HAZ-scores measure the cumulative investments that parents have made in children over time and are particularly sensitive to investments made while children are young (World Health Organization 1983). BMIZ and WAZ-scores fluctuate more contemporaneously and can therefore highlight short-term changes in nutrition. In addition, we focus on young children aged 0 to 5 in the initial survey years since the potential effects on young children can be long lasting when their parents migrate (Parreñas 2005); moreover, we improve the identification by constructing instrumental variables for changes in parental migration status. The instrumental variables are the interaction terms of wage growth in provincial capital cities with initial village migrant networks. Finally, we provide suggestive evidence about whether parent migration affects housing conditions and hygiene in the living environment, which can affect children's nutrition.

Our results show that parental migration has no significant impact on young children's HAZ-scores, but it improves their WAZ-scores. We provide suggestive evidence that the improvement in short-term nutrition status may be achieved through increased access to tap water among households with migrants.

The paper proceeds as follows. Section 2 provides more background about China's rural-to-urban migration and nutritional status among children in rural areas. The third section presents a conceptual framework, and the fourth section discusses the data we use and the empirical framework for the paper. The econometric results are presented and discussed in the fifth section. The final section concludes.

2. BACKGROUND

Accelerating rural-to-urban migration has been an important demographic feature in China during the past three decades. In this paper, we are particularly interested in the implications of migration for the nutritional status of children left behind in rural villages. In this section, we will outline the evolution of China's internal migration and its implications for rural communities, followed by a discussion about how the distribution of children's nutritional status is changing in rural China.

Migration and Children Left Behind in Rural China

China's labor market experienced dramatic changes during the 1990s as the number of rural residents moving to urban areas for employment grew rapidly. Estimates using the 1 percent sample from the 1990 and 2000 rounds of the Population Census and the 1995 1 percent population survey suggest that the intercountry migrant population grew from slightly more than 20 million in 1990, to 45 million in 1995, to 79 million by 2000 (Liang and Ma 2004). These figures likely underreport the true scale of migration because they do not account for migration that takes place during shorter periods of time (Cai, Park, and Zhao 2008).

The expansion of migration during the past two decades has been facilitated by the relaxation of constraints on the household registration (*hukou*) system. The initial hukou reform, taking place in 1988, established a mechanism for rural migrants to obtain legal temporary residence in China's urban areas (Mallee 1995). After rural migrants could obtain legal temporary residence, they became better able to establish networks to facilitate the job search in distant labor markets (for example, Munshi 2005); however, legal temporary resident status does not imply access to urban health and education services or social safety nets because these benefits are still linked to household registration status (World Bank 2009). This institutional arrangement effectively discourages entire families from moving to cities. Whereas individuals with rural hukou status can now purchase nonagricultural hukou status from urban governments in many places, the system continues to work against more permanent migration flow (Fan 2008). Migrants consequently maintain close ties with their ancestral villages.

The impacts of migration on rural households and individuals left behind are mixed. On one hand, remittances benefit migrant-sending households, as evidenced by their higher consumption (Du et al. 2005; de Brauw and Giles 2008), improved risk-coping ability (Giles 2006; Giles and Yoo 2007), and higher levels of investment in productive assets (Woodruff and Zenteno 2007). On the other hand, individual household members who are left behind, particularly women and seniors, may be somewhat negatively affected. For example, women who live in migrant households now perform more farm work than they would have otherwise (Mu and van de Walle 2011), and the elderly with migrant adult children are at greater risk of falling into poverty (Giles, Wang, and Zhao 2011). Even though children's educational performance is not compromised but actually improves in migrant households (Chen et al. 2009), some nutrition measures of school-age children are negatively affected (de Brauw and Mu 2011).

It is important to remember that in the mid-1990s, migration rates were substantially higher among men than women for most age ranges (Rozelle et al. 1999). Since the late 1990s more women have participated in rural-to-urban migration, particularly the young and the single (Du, Park, and Wang 2005; de Brauw et al. 2008). In some provinces, the gender gap in migration rates is still increasing (Mu and van de Walle 2011). Consistent with differences in migration prevalence by gender shown elsewhere, we demonstrate later in the paper that the majority of parental migration is by fathers.

Nutritional Status of Rural Children

China faces multiple challenges in improving nutrition among rural children. First, disparities are present in child health in rural and urban areas as well as within and among regions. For example, under nutrition in western provinces is acute, and stunting and underweight are three times more prevalent in rural than in urban areas (UNICEF 2008). Child growth is becoming more sensitive to household income, a consequence of increasing income inequality and the disappearance of subsidized food coupons in the mid-1990s (Osberg, Shao, and Xu 2009). In addition, a significant share of rural children remains malnourished. In 2002, the nationwide stunting rate was still nearly 15 percent (Svedberg 2006). With inadequate access to regular micronutrient-rich diets, many rural children also suffer from iron deficiency (Luo et al. forthcoming). At the same time, induced by increased income, the structure of the Chinese diet is shifting away from high-carbohydrate foods toward high-fat, high-energy-density foods (Du et al. 2004); consequently, overweight status has been increasing among children in both rural and urban areas (Monda and Popkin 2005; de Brauw and Mu 2011).

Notwithstanding these challenges, the average measures of the nutritional status of children have clearly improved over time. Economic growth is well known to be associated with improved nutritional status (Fogel 2004), and China's experience is no exception. Child growth improved during the early reform period. In 1990, the average height of children aged two to five in seven provinces was 92.5 cm (centimeters) in rural areas, an increase of 3.8 cm from the average in 1975 (Shen et al. 1996). Improvement in children's growth has continued since then as children's average HAZ, WAZ, and weight-for-height *z* (WHZ) scores increased dramatically between 1990 and 2000 (Chen 2000; Osberg, Shao, and Xu 2009). These increases represent an anthropometric confirmation of significant progress in average well-being.

Some improvements in child growth are directly attributable to the progress made in public service provision. A salient example is water delivery because unsafe drinking water can cause diseases such as diarrhea, which can severely affect child growth (for example, Glewwe and Miguel 2008). Efforts to establish public water systems in urban China date back to the 1950s. During the 1980s, the Chinese government launched a drinking water improvement program in rural areas (Zhang 2012). Clean water from water plants improves rural children's heights and weight-for-height *z* scores substantially (Zhang 2012). In the same context, Mangyo (2008) also finds that access to in-yard water sources improves child health when mothers are relatively well educated. Even though the placement of village-level water projects may be exogenous to individual decisions, household-level water access may still be affected by individual demand (Mangyo 2008). Motivated by these findings, in examining the channels through which migration may affect children's nutrition, we will explicitly investigate whether parental migration improves household access to clean water.

3. CONCEPTUAL FRAMEWORK

One can hypothesize four main channels through which migration might affect nutritional investment in children: an income effect, an information effect, a time effect, and an effect from changes in the structure of intra-household bargaining and cooperation.

First, from the perspective of the household, a primary reason to send out migrants is to increase overall household income, no matter how migration or the household is defined.4 An increase in household income can lead to better nutritional outcomes among children through various means. For example, higher incomes could lead to increased calorie intake if food is scarce within the household (Fogel 2004). The demand elasticity's of nutritious foods, however, tend to be higher than those of coarse grains (Bouis 1994); therefore, households may use income increases to purchase more foods rich in micronutrients (such as fruits and vegetables) and protein (such as eggs and meat) (Subramanian and Deaton 1996; Nguyen and Winters 2011). Improvements in diet might manifest themselves particularly in HAZ scores. In addition, higher income could also lead to improved housing conditions and a more hygienic environment, such as an environment with better access to clean water. Finally, health service utilization for children may increase with income. The majority of rural residents were uninsured before the New Cooperative Medical Scheme was introduced in 2003, but even since 2003, out-of-pocket medical spending still positively correlates with income (Wagstaff et al. 2009).

Second, independent of income effects, migration experience may improve household health information and hygiene practices. As a result, the demand for healthy food, medical service, clean water, and better sanitation may increase in migrant households. Similar to income effects, information effects are expected to improve nutrition outcomes. We cannot separate these effects in this study, but it is worth noting that either of them can bring a positive impact as a result of migration.

Third, by sending out migrants, households decrease their overall labor endowment. Because households in rural China are almost always engaged in some form of local income generation, either through farming or self-employment, child care time becomes scarcer in migrant-sending households. As a result, less time is allocated to cooking (Chen 2006; de Brauw and Mu 2011) and monitoring the eating habits of children; therefore, to the extent that it is difficult and time-consuming to ensure nutritional investments in children, the time effect of migration may compromise child nutrition, especially among young children.

Fourth, when household members live apart as a consequence of migration, the equilibrium in the implicit intra-household bargain may change. For example, some argue that wives gain increased autonomy and new decision making powers as household heads after their husbands migrate (Davin 1999). Considering that the empowerment of women is generally associated with better child health outcomes (for example, Thomas 1990; Duflo 2003), one may expect that fathers' migration could positively affect child nutrition. Alternatively, the adult household members left behind may use more child labor in household chores but counter any potential negative effect on children by increasing their food intake. Such strategic behavior may result in an innocuous impact on child nutrition (Chen 2006).

As outlined in this conceptual framework and suggested by mixed results from existing studies, whether parental migration affects children's health and nutrition is primarily an empirical question. In the next section we present the data and empirical framework we will use to try to measure this relationship in rural China.

[4] To be precise, we consider the migrant still a household member after migrating, at least until he or she sets up a household at the destination. So long as the returns to human capital at the destination exceed those at the source, full household income would increase. Among those left behind, household income increases if migrant remittances exceed total household income without the migrant's contribution.

4. DATA AND EMPIRICAL FRAMEWORK

Data

The data used in the analysis derive from the four waves of the China Health and Nutrition Survey (CHNS) of 1997, 2000, 2004, and 2006.5 The CHNS is a longitudinal survey covering nine provinces that vary substantially in geography, economic development, and access to public resources.6 The survey includes the demographic characteristics of each household as well as asset holdings. We further make use of community-level questionnaires that accompanied the household survey.

Important for this paper, the CHNS records anthropometric measurements, height, and weight for each individual within the household. To construct HAZ, WAZ, and BMIZ scores, we use the most recent growth charts made available by the World Health Organization (WHO), which have corrected previous standard growth charts for what are now considered developed-country biases (de Onis et al. 2007).[7]

The CHNS was not originally conceived to study migration, so the migration status of household members in the CHNS must be built up from the household roster. Starting from 1997, if an individual who was in a previous round of the CHNS is not in the current round of the survey, a question is asked regarding why this individual does not currently reside in the household. We consider any individual who has left the home to seek employment between any two waves of the survey to be a migrant.[8]

We pool panel data from all three pairs of surveys—1997 and 2000, 2000 and 2004, and 2004 and 2006—keeping in the sample children younger than five years in the initial years (1997, 2000, and 2004), and survey for two consecutive rounds in our analysis. Our focus on young children is motivated by the findings that they are at greater risk for problems associated with malnutrition and are more likely to respond to nutrition interventions (WHO, 1983). With two observations for each child, we can apply panel data analysis to control for time-invariant individual, family, and community characteristics.

Trends related to migration and anthropometric measures found in the literature, discussed in section 2, are also present in the CHNS. To illustrate migration trends, we graph locally weighted regressions of out-migration on age for individuals aged 16 to 75 (see Figure 4.1). Age is negatively correlated with the probability of out-migration for both men and women. Except for the oldest age cohorts, migration prevalence among individuals of any given age increases over time among both men and women. The rate of increase is slower among older individuals, implying that migrant labor flows tend to be young and getting younger, consistent with findings in other studies (for example, Liang and Ma 2004). In the 25- to 40-year-old range, migration rates tend to increase among younger workers, more likely to be married and with children, from around 10 percent for men in 1997 to between 30 and 35 percent in 2006. For women, increases in migration are less dramatic, from less than 10 percent in 1997 to greater than 20 percent in 2006. It could reasonably be expected that migrations by individuals in these cohorts have the largest chances of affecting the nutritional status of children.

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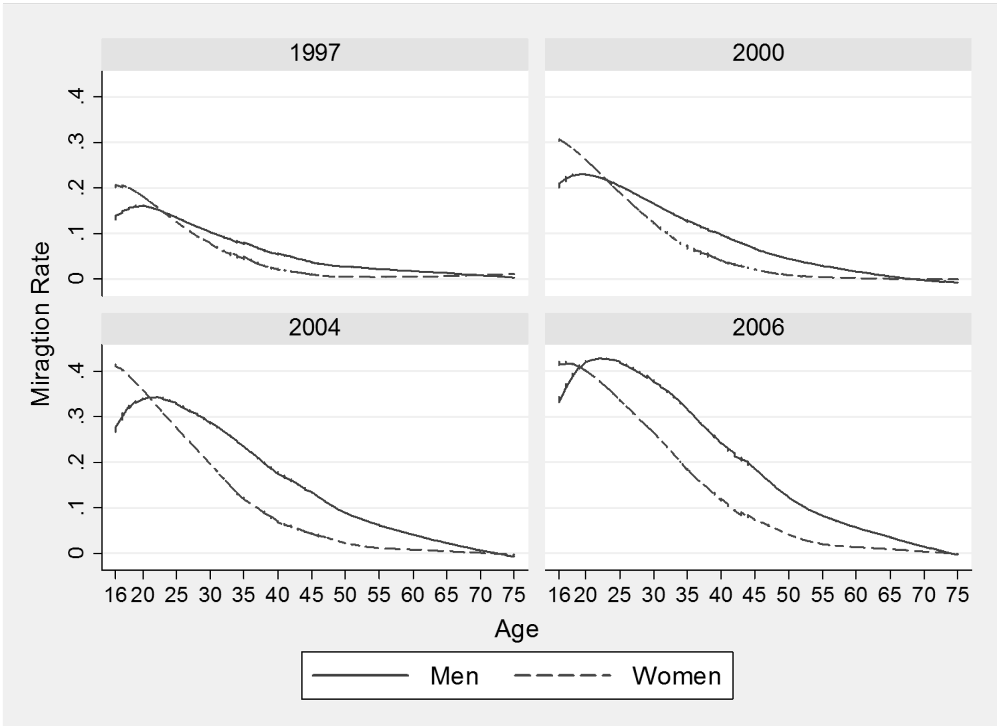
[5] See http://www.cpc.unc.edu/projects/china for details.

[6] The nine provinces are Henan, Hubei, Heilongjiang, Liaoning, Shandong, Guizhou, Jiangsu, Guangxi, and Hunan.

[7] See http://www.who.int/childgrowth/en/ and http://www.who.int/growthref/en/ for details about the growth reference data, which are consistent with one another.

[8] This definition differentiates migration for work from migration for other reasons, but it could underestimate the migration rate because return migrants cannot be identified, nor can temporary migrants. Hence, we may underestimate the impact of parents' migration on children's nutrition status.

Figure 4.1—Migration rate, by gender and year, China, 1997-2006



Source: China Health and Nutrition Survey, 1997–2006.

Consistent with the evidence presented in Figure 4.1, over time, more children in our sample are found to live in migrant households (see Table 4.1). Whereas 6 percent of children had a migrant parent in 1997, by 2006 this figure increased to nearly 34 percent. Most parental migration in China is done by fathers: Migration by the father alone increases from 3.4 percent of households in 1997 to 13.3 percent in 2006. In later rounds, it is increasingly common to observe both parents' migrating—this case is found among 10 percent of children in 2004 and among 15.7 percent of children in 2006. For a small proportion of children, the mother is the sole migrant.

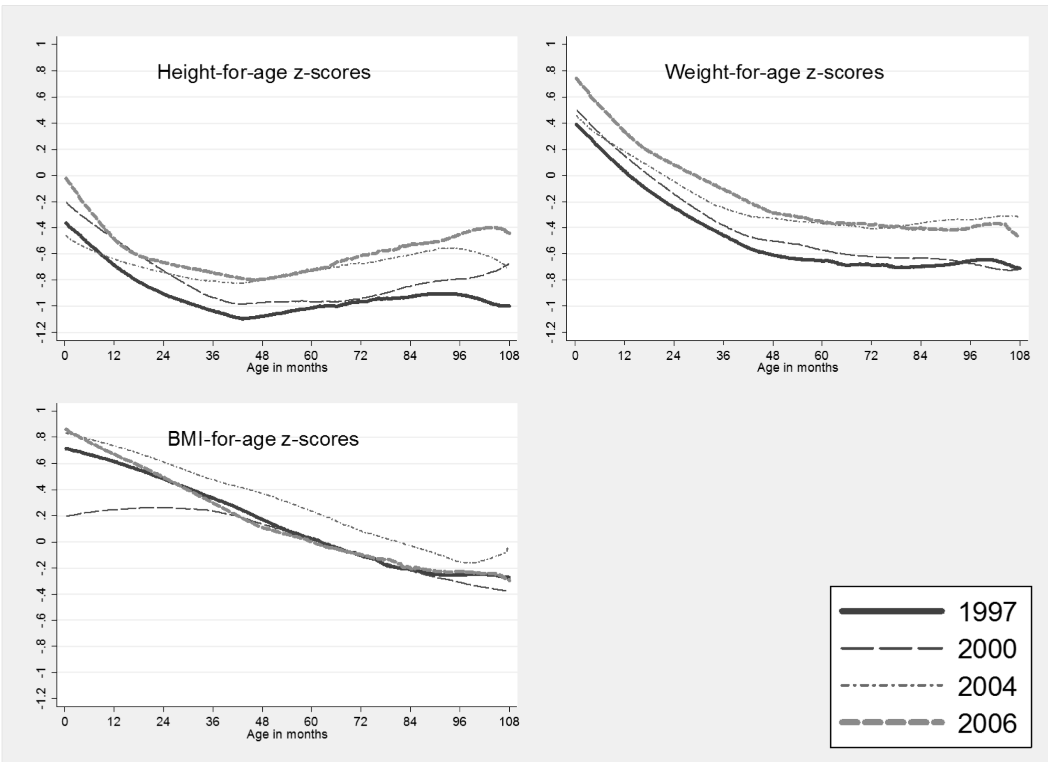
Table 4.1—Migration of parents of rural children, China, 1997-2006

	1997 (aged 0–5)	2000 (aged 0–8)	2004 (aged 0–9)	2006 (aged 2–7)
Any parent migrated	0.060	0.116	0.231	0.337
	(0.238)	(0.320)	(0.422)	(0.476)
Only father migrated	0.034	0.069	0.090	0.133
	(0.181)	(0.254)	(0.286)	(0.341)
Only mother migrated	0.009	0.019	0.040	0.047
	(0.092)	(0.136)	(0.197)	(0.213)
Both parents migrated	0.017	0.027	0.100	0.157
	(0.130)	(0.163)	(0.300)	(0.366)
Number of observations	239	479	325	85

Source: China Health and Nutrition Survey, 1997-2006
Note: Standard deviations are in parenthesis. Rural children who were aged zero to five in 1997, 2000, or 2004 and surveyed in the China Health and Nutrition Survey for at least two consecutive rounds with nonmissing anthropometric measures are included in the calculation.

In Figure 4.2, we present locally weighted regressions of HAZ, WAZ, and BMIZ for children younger than 10. On average, HAZ and WAZ tend to increase over time at most ages. In 1997, average HAZ scores are less than –1, implying a significant stunting rate.[9] By 2006, average HAZ increase to around –0.7, with the scores of very young children approaching those of the healthy reference population. The age profile of average HAZ scores shows that this measurement deteriorates from birth to approximately 48 months. After that, average HAZ scores level off or improve slightly. This pattern is similar to patterns found in Guatemala (Carletto et al. 2011), where the turning point is around 30 months. WAZ scores are negatively correlated with age, a pattern common in many developing countries, such as El Salvador (de Brauw 2011). A steady increase in WAZ scores is evident during the period between 1997 and 2006. Children younger than 3 have already had average WAZ scores equal to or greater than, the international norm in 2006. Given that WAZ scores for most children in the sample are higher than their HAZ scores, it is not surprising that BMIZ scores are actually higher than international norms on average until children reach age 5.

Figure 4.2—Nutritional outcomes of children aged zero to nine, by year, China, 1997-2006

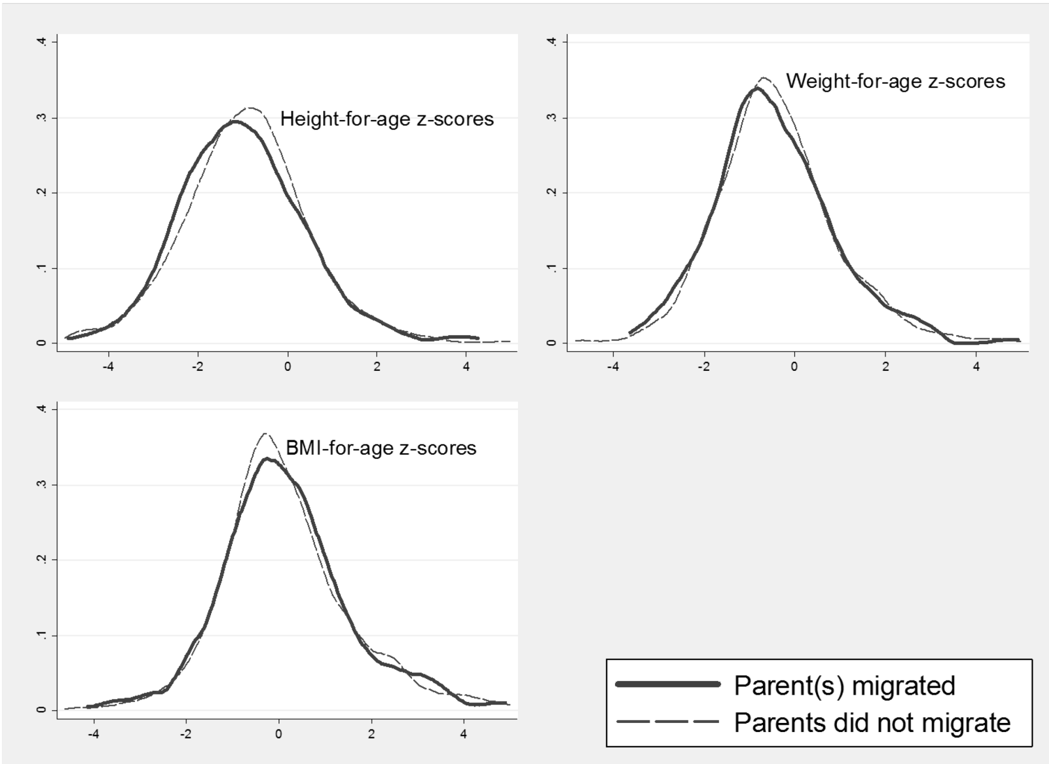


Source: China Health and Nutrition Survey, 1997–2006

We explore the differences in distribution of these anthropometric measures between children with at least one migrant parent and those whose parents who did not migrate at the time of the survey. Figure 4.3 shows that the middle of the distribution of HAZ scores for children whose parents did not migrate is positioned slightly right of that for children with a migrant parent, but no distinct shift in the whole distribution is observed. The distributions of WAZ and BMIZ scores are similar for these two groups of children.

[9] Children with HAZ scores less than –2 are considered stunted.

Figure 4.3—Nutritional outcomes of children aged zero to nine, by parental migration status, China, 1997-2006



Source: China Health and Nutrition Survey, 1997–2006

Empirical Framework

The conceptual framework in section 3 suggests estimating an equation for nutritional outcomes of individual child *i* in household *h* of village *j* at time *t* denoted by the nutrition outcome *H* of child i in household h located in village j at time *t* (H_{ihjt}), which is a function of parents' migration status (M_{hjt}) measured by a dummy variable equal to 1 if at least one parent has migrated out at time *t,* and 0 otherwise; the age and gender of the child at the initial years (A_{ihj0}); characteristics of the household (X_{hjt}) and the village (V_{jt}); province-specific year effects $Y_{p\times t}$; and individual fixed effects (u_i). To allow age and gender effects to vary over time to account for the way children grow as they age in the early years, we interact A_{ihj0} with a time trend *T.* With the idiosyncratic error term denoted as ε_{ihjt}, the estimation equation can be written as follows:

$$H_{ihjt} = \alpha_0 + \alpha_1 M_{hjt} + A_{ihj0} T\alpha_2 + X_{hjt}\alpha_3 + V_{jt}\alpha_4 + Y_{p\times t} + \mu_i + \varepsilon_{ihct} \quad (1)$$

Our conceptual framework suggests that controlling for heterogeneity between households and villages is likely to be important in explaining nutritional outcomes among children. To this end, the rich information in the CHNS is useful because it allows us to include a large set of control variables in the regression analysis. Specifically, the X_{hjt} vector includes the years of education of the household head, the number of senior (aged 60 and older) and working-age (aged 16 to 59) household members by gender, household size, and household assets per capita. We include an indicator for the presence of a health clinic in the village and the share of other households in the village that have tap water in the vector of village characteristics (V_{jt}). We also include the village-level market prices of rice, flour, and pork because prices may affect both relative wealth and household diet and therefore may influence both migration decisions and the nutritional status of children. Finally, the province–time fixed effects ($Y_{p\times t}$) control for time-varying macroeconomic conditions that differ by provinces.

With this large set of control variables, we still cannot rule out that unobservable traits of a child, such as innate health, can affect both the migration decision of their parents and their growth outcomes. With a panel of children, a reasonable way of controlling for individual-level fixed effects is to difference the two periods. After differencing equation (1), we can write the following:

$$\Delta H_{ihjt} = \alpha_1 \Delta M_{hjt} + A_{ihj0}\alpha_2 + \Delta X_{hjt}\alpha_3 + \Delta V_{jt}\alpha_4 + \Delta Y_{p\times t} + \Delta \varepsilon_{ihct}. \quad (2)$$

In estimating equation (2), any effect of migration on children's nutritional outcomes is identified off changes in migration status among households occurring between survey rounds; however, we still must be concerned that unobservable that vary over time exist and both are important determinants of child nutritional status and affect household migration decisions. Clearly, such variables are not accounted for in equation (2). For example, income shocks resulting from crop failures could affect both migration decisions and the diets of children (or other factors that influence their nutritional status). To identify the impact of migration, we must find variables that affect migration but do not independently affect children's nutritional status, except through their effects on migration. In this context, we suggest the following identification strategy. We hypothesize that relative wage growth in provincial capital cities (ΔW_{pt}) potentially affects migration decisions by households. We further hypothesize that the strength of that signal will be affected by the size of the village migrant network or the initial size of the village migrant network, M_{-hj0}, as measured in the first available survey. We specifically measure the migrant network measured as the share of men and women in the working-age population who had migrated in the initial year.[10] The change in migration can therefore be written as follows:

$$\Delta M_{hjt} = f(\Delta W_{pt}) + rM_{-hj0} + M_{-hj0} \times f(\Delta W_{pt}). \quad (3)$$

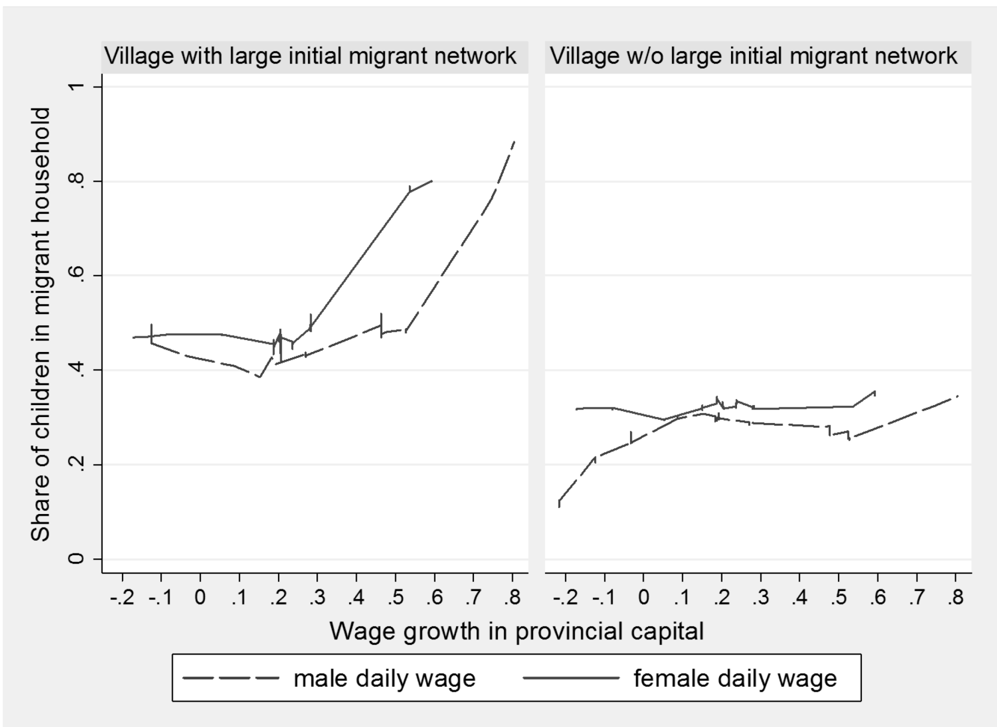
The instrumental variables strategy then follows. We hypothesize that wage growth and the migrant network individually affect migration decisions; however, on their own, each variable also might be correlated with children's nutrition outcomes even after we account for migration. Wage growth in the provincial capital might be correlated with local wage growth, which could be reflected in changes in anthropometric measures through the income effect. Similarly, households with large migrant networks might have better access to information on nutrition or hygiene practices that affect anthropometric outcomes independently of migration; however, we argue that the interaction of the migrant network variable and provincial capital wage growth will affect only migration, and once we control for the migration network and the provincial capital wage growth directly, the variation in the interaction term will affect migration and will not affect independently children's anthropometric outcomes. We note that in equation (2), we implicitly control both for wage growth ΔW_{pt}, through the province-specific year effect ($\Delta Y_{p\times t}$), and for M_{-hj0} as it is included in the fixed effect and is therefore differenced out. The interaction term alone ($\Delta W_{pt} \times M_{-hj0}$) is therefore used for statistically identifying migration.

Intuitively, if we control for wage growth, migration is more likely among individuals with larger migrant networks, and therefore their children are more likely to be left behind. We confirm that this relationship exists in the data in Figure 4.4, in which we stratify villages in the sample into two groups: one with larger migrant networks in the initial year and a group with smaller migrant networks in the initial year. We define the cutoff for large migrant networks as villages in which 15 percent of working-age women or 20 percent of working-age men migrated in 1997. The graphs illustrate that in villages with

[10] The reasons for absence were recorded only from the 1997 round on, so we cannot use earlier waves to construct the instrumental variables. For villages in Liaoning province, which were added to the survey in 2000, the 2000 migration rates are used to measure the migrant network.

larger migrant networks, migration grows faster with wage growth but that the same relationship is weaker in villages with smaller migrant networks. Moreover, this pattern holds for both men and women. This approach essentially exploits the complementarily between urban wage and migrant network in an individual's migration decision. Under the assumption that the interaction variable is independent of $\Delta\varepsilon_{ihct}$ in equation (2), our estimate of α_1 would be unbiased. We primarily test this assumption by checking the robustness of the estimates to the inclusion of various sets of variables that might confound the relationship between migration and the interaction variable.

Figure 4.4—Share of children in migrant households and wage growth in provincial capital, by village initial migrant network, China, 1997-2006



Source: China Health and Nutrition Survey, 1997–2006.
Note: w/o = without. Villages with large initial migrant networks are defined as villages in which 15 percent of working-age women or 20 percent working-age men migrated in initial years. The rest of the villages are labeled as villages without large initial migrant networks.

5. RESULTS

Primary Results

We first examine HAZ scores as the dependent variable and estimate equations (1) and (2), respectively, treating parental migration as exogenous. The results are reported in Table 5.1. In the ordinary least squares regression, household head's years of schooling, the number of working-age women, household assets, and the percentage of other households in the village having tap water are shown to be positively correlated to children's HAZ, but these correlations cease to be significant in the first differencing estimation. Two estimations yield small coefficients with large standard errors on the measure of migration, which is an indicator variable for whether a parent of the child migrated. The results seem to suggest that parental migration status may not affect children's cumulative nutrition. But these point estimates are likely to be biased because the endogeneity of migration is not accounted for.

Table 5.1—Baseline regression results on children's height-for-age *z* scores, China, 1997-2006

Dependent Variable: Height-for-age z Scores	**Ordinary Least Squares**	**First *Differencing***
Parent(s) migrated	0.006	–0.078
	(0.178)	(0.179)
Individual Characteristics		
Boy × Time Trend	0.043*	–0.039
	(0.022)	(0.031)
One Year Old × Time Trend	–0.008	0.107
	(0.036)	(0.094)
Two Years Old × Time Trend	0.057	0.237**
	(0.036)	(0.100)
Three Years Old × Time Trend	0.057**	0.185*
	(0.029)	(0.101)
Four Years Old × Time Trend	0.005	0.188*
	(0.038)	(0.101)
Five Years Old × Time Trend	0.056	0.544*
	(0.177)	(0.291)
Household Characteristics		
Head's years of schooling	0.037**	0.014
	(0.018)	(0.016)
Number of females aged 60 and above	0.335**	0.058
	(0.153)	(0.186)
Number of males aged 60 and above	–0.012	0.227
	(0.191)	(0.239)
Number of working-age men	–0.081	0.098
	(0.103)	(0.105)
Number of working-age women	0.297***	–0.070
	(0.105)	(0.130)
Household assets per capita (log)	0.128***	0.025
	(0.041)	(0.035)
Village Characteristics and Food Prices		
Free market price of pork	0.019	–0.021
	(0.019)	(0.021)
Free market price of regular rice	–0.010	–0.003
	(0.034)	(0.025)
Free market price of flour	–0.009	0.032*
	(0.019)	(0.019)
Village has a clinic	–0.029	–0.125
	(0.145)	(0.162)
Other households with tap water (%)	0.318**	–0.047
	(0.159)	(0.187)
Province-specific year dummies	yes	yes
Number of observations	935	491

Source: China Health and Nutrition Survey, 1997–2006.
Notes: Standard errors in parenthesis are robust to heteroscedasticity and permit within-village correlations in unobservable. *Significant at 10 percent. **Significant at 5 percent. ***Significant at 1 percent.

Next, we use the interaction terms described above as instrument variables for migration, and we re-estimate equation (2) using an instrumental variables estimator (see Table 5.2). The instruments are gender-specific interactions between growth in wages in the provincial capital and the initial village migration rate. With the full set of control variables for age and gender of the child, household characteristics, and village characteristics, we find that the instruments are strongly correlated with changes in migration status among parents (column 1); both have *z* statistics between 2 and 3, and the *F* statistic on the joint hypothesis test that both coefficients are zero is 7.8. The under identification test and the over identification test are passed comfortably, but under standard assumptions about the error terms, the instruments would be considered weak (for example, Stock and Yogo 2007); however, in the presence of clustering, it is less clear whether an *F* statistic between 7 and 8 is associated with weak instruments (for example, Finlay and Magnusson 2009). In the second stage, we again find no significant relationship between the migration variable and child HAZ (see Table 5.2, column 2). As expected, the standard errors rise whereas the point estimate remains small. Consistent with results reported in Table 5.1, these results confirm that parental migration does not affect children's HAZ scores.

Table 5.2—Instrumental variables estimates of the determinants of children's height-for-age z scores: China, 1997-2006

Dependent Variable: ∆Height-for-age z Scores	**First Stage**	**Second Stage**
∆Parent(s) migrated		0.056
		(0.738)
Individual characteristics	yes	yes
Household characteristics	yes	yes
Village characteristics and food prices	yes	yes
Province-specific year dummies	yes	yes
Instrument variables		
Growth of Male Daily Wage in Provincial	0.112**	
Capital City × Initial Village Male Migration Rate	(0.047)	
Growth of Female Daily Wage in Provincial	0.119**	
Capital City × Initial Village Female Migration Rate	(0.053)	
F test on excluded instruments	7.865	
Probability > *F*	0.001	
Under identification: Kleibergen-Paap rk LM statistic	11.48	
Chi-square(2) p-value	0.003	
Over identification: Hansen J statistic	0.109	
Chi-square(1) p-value	0.741	
Weak identification: Cragg-Donald Wald *F* statistic	8.157	
Number of observations	491	

Source: China Health and Nutrition Survey, 1997-2006.
Notes: Standard errors in parenthesis are robust to heteroscedasticity and permit within-village correlations in un observables.
**Significant at 5 percent.

Although migration does not affect the cumulative measure of nutrition, it may affect shorter-term measures of nutrition status, such as BMIZ and WAZ scores. The estimated results of these variables based on equation (2) with the instrument variable method are reported in Table 5.3. We find a point estimate of 0.102 on parental migration status for BMIZ, but it is not statistically different from zero. When we estimate the model using changes in WAZ scores as the dependent variable, we find statistically significant coefficients of a reasonable magnitude: 0.201. This estimate essentially suggests that the children of migrants are heavier than the children of others, ceteris paribus, at the same age among children of the same gender. Given the absolute lack of significance when using HAZ scores as the dependent variable, migration does not appear to affect height; so these estimates lend credence to the ideas that BMIs are increasing and that impact is on short-term instead of long-term nutritional measures. This line of argument would suggest that our inability to reject the null hypothesis that the coefficient estimate of the BMIZ scores is different from zero is the result of a lack of statistical power in the regression.

Table 5.3—Instrumental variables estimates of the determinants of children's BMI-for-age z scores (BMIZ) and weight-for-age z scores (WAZ), China, 1997-2006

	∆BMIZ	∆WAZ
∆Parent(s) migrated	0.102	0.201**
	(0.067)	(0.079)
Individual characteristics	yes	yes
Household characteristics	yes	yes
Village characteristics and food prices	yes	yes
Province-specific year dummies	yes	yes
F test on excluded instruments	7.879	8.093
Probability > *F*	0.001	0.000
Under identification: Kleibergen-Paap rk LM statistic	13.88	12.01
Chi-square(2) p-value	0.000	0.002
Over identification: Hansen J statistic	1.406	0.306
Chi-square(1) p-value	0.236	0.580
Weak identification: Cragg-Donald Wald *F* statistic	9.065	7.266
Number of observations	455	511

Source: China Health and Nutrition Survey, 1997–2006.
Notes: Standard errors in parenthesis are robust to heteroscedasticity and permit within-village correlations in unobservable.
**Significant at 5 percent.

In summary, our primary regression results indicate that parental migration has a positive effect on children's weight but has no impact on height. Next we check the robustness of our estimates to potential misspecifications and confounding factors. We will then explore the mechanisms by which this effect may have taken place.

Robustness Checks

One of the primary concerns about our results is that changes in household characteristics may be endogenous to children's nutritional status. For example, the change in household assets may be correlated with unobserved shocks that could also affect changes in children's nutritional outcomes; moreover, asset changes or changes in household demographic compositions may be correlated with the migration decisions of household members. Such correlation may lead to biased estimates of migration. To rule out the possibility that the above results are driven by changes in endogenous household characteristics, we estimate the regressions without including household characteristics as control variables. The results are reported in Table 5.4.

Compared to the previous estimates with household-level control variables, the coefficients on migration in Table 5.4 are smaller in all three regressions. One possible reason for such difference is that parental emigration negatively affects the number of household members, whereas the number of household members, female members in particular, seems to be positively correlated with children's nutrition outcomes.[11] Albeit the change in the magnitude, the results are consistent with the previous finding that parental emigration positively affects children's weight but not their height.

[11] The number of female household members aged 60 and older and the number of working-age women have positive and significant coefficients in the BMIZ and WAZ score regressions. These results are not reported in Table 5.3, but they are available upon request.

Table 5.4—Robustness check—instrumental variables estimations without household characteristics, China, 1997-2006

	ΔHeight-for-age z Scores	ΔBMI-for-age z Scores	ΔWeight-for-age z Scores
ΔParent(s) migrated	0.019	0.030	0.082**
	(0.421)	(0.042)	(0.036)
Individual characteristics	yes	yes	yes
Household characteristics	no	no	no
Village characteristics and food prices	yes	yes	yes
Province-specific year dummies	yes	yes	yes
F test on excluded instruments	11.78	7.879	8.093
Probability > *F*	0.000	0.001	0.000
Underidentification: Kleibergen-Paap rk LM statistic	12.50	13.88	12.01
Chi-square(2) p-value	0.002	0.000	0.002
Over identification: Hansen J statistic	0.057	1.406	0.306
Chi-square(1) p-value	0.812	0.236	0.580
Weak identification: Cragg-Donald Wald *F* statistic	16.71	9.065	7.266
Number of observations	491	455	511

Source: China Health and Nutrition Survey, 1997–2006.
Notes: Standard errors in parenthesis are robust to heteroscedasticity and permit within-village correlations in unobservable. **Significant at 5 percent.

We are also concerned about potential panel attrition bias. The CHNS sample attrition is high for children aged zero to five, particularly between 2004 and 2006 (see Table 5.5). If the attrition is systematically correlated with children's nutrition status and parental migration decision, then the estimation based on the remaining panel would be biased. In Table 5.5 we compare these variables between children who were surveyed in two consecutive rounds with those who dropped after the initial years, and we find neither statistically significant differences in average anthropometric scores measured in the initial years nor differences in the propensity to come from a migrant household. For this set of observables, panel attrition would therefore appear to be random.

Table 5.5—Anthropometric measures among children and parent migration status, by panel attrition, China, 1997-2006

	Mean Difference by Attrition		
	Stayers (1)	Leavers (2)	(1)-(2) (3)
1997–2000 panel			
Height-for-age *z* score	–1.159	–0.937	–1.445
Weight-for-age *z* score	–0.524	–0.253	–1.812
BMI-for-age *z* score	0.371	0.535	–1.037
Parent(s) migrated	0.064	0.073	–0.315
Number of children	219	116	
2000–2004 panel			
Height-for-age *z* score	–1.218	–1.163	–0.321
Weight-for-age *z* score	–0.452	–0.308	–0.907
BMI-for-age *z* score	0.359	0.304	0.323
Parent(s) migrated	0.132	0.177	–1.094
Number of children	219	133	
2004–2006 panel			
Height-for-age *z* score	–1.152	–0.993	–0.678
Weight-for-age *z* score	–0.214	–0.096	–0.586
BMI-for-age *z* score	0.359	0.601	–1.189
Parent(s) migrated	0.233	0.259	–0.454
Number of children	73	220	

Source: China Health and Nutrition Survey, 1997–2006.
Note: Sample means are reported in columns 1 and 2. The last column reports t statistics for testing the equality of the two sample means.

Nonetheless, we re-estimated the main IV regressions using a correction for attrition as follows. First, we estimated the probability that individuals stayed in the sample, given initial characteristics of individuals (Table A.2). We subsequently use the inverse of that estimated probability as a weight in the regressions (Wooldridge 2010). The intuition is that by giving more weight to children who appear less likely to remain in the sample, we make the analysis sample more representative of the original one. Continuing to control for the full set of other explanatory variables (see Table 5.6), we find that the results are largely unchanged. We find statistically significant estimates for the effect of migration on WAZ scores but not on BMIZ or HAZ scores.[12]

Table 5.6—Robustness check—instrumental variables estimations with inverse probability weighting to correct for panel attrition, China, 1997-2006

	∆Height-for-age z Scores	∆BMI-for-age z Scores	∆Weight-for-age z Scores
∆Parent(s) migrated	–0.367	0.085	0.202**
	(0.872)	(0.064)	(0.081)
Individual characteristics	yes	yes	yes
Household characteristics	yes	yes	yes
Village characteristics and food prices	yes	yes	yes
Province-specific year dummies	yes	yes	yes
F test on excluded instruments	6.110	8.326	8.471
Probability > F	0.003	0.000	0.000
Under identification: Kleibergen-Paap rk LM statistic	12.14	14.21	13.24
Chi-square(2) p-value	0.002	0.000	0.001
Over identification: Hansen J statistic	0.001	2.514	0.303
Chi-square(1) p-value	0.973	0.113	0.582
Weak identification: Cragg-Donald Wald F statistic	8.400	9.597	7.624
Number of observations	491	455	511

Source: China Health and Nutrition Survey, 1997–2006.
Notes: Standard errors in parenthesis are robust to heteroscedasticity and permit within-village correlations in unobservable. **Significant at 5 percent.

Mechanisms

Given the pattern of results, we want to understand how migration might have affected children's weight. As discussed in section 3, weight and height could be affected by an increase in the quantity and quality of food consumed, respectively. When we estimate the effects of migration on caloric intake in a framework outlined in equation 2, we find no evidence that either the quantity or the quality of young children's diet has changed because of the migration of their parents (see Table 5.7).[13]

A second potential mechanism for increased weight discussed in section 3 is through improved hygiene. To measure whether various measures of hygiene could explain our findings, we estimate equation (2) using three measures of the children's living environment as the dependent variable (Table 5.7). We find no statistically significant relationship between migration and having a flush toilet in the house or having any excreta around the house. We do, however, find that households with migrants are 7 percentage points more likely to have a water tap in the house or in the yard by the second survey relative to non-migrants. Finally, we find no relationship between child healthcare variables, namely, immunization and preventive health service, and parental migration. So it seems that the plausible mechanism we can identify is through increases in the availability of tap water.

[12] We are not able to precisely estimate heterogeneous impacts along the gender or poverty dimensions.

[13] However, it is worth noting a problem with this approach, which is that demand for calories increases among children as they age, so it is likely difficult to measure this effect even if it existed. We do not report results on the share of calories consumed from carbohydrates, protein and fat, for reasons of brevity; estimates are available from the authors.

Table 5.7—Instrument variables estimates of impacts of migration on hygiene environment and health inputs, China, 1997-2006

	ΔCalorie Intake (Kilocalorie per Day [Log])	ΔIn-house Flush Toilet	ΔNo Excreta around House	ΔIn-house or In-yard Tap Water	ΔImmunization Last Year	ΔPhysical Checkup Past Four Weeks
ΔParent(s) migrated	0.418	0.458	0.113	0.073**	0.106	–0.148
	(0.349)	(0.392)	(0.580)	(0.031)	(0.317)	(0.120)
Individual characteristics	yes	yes	yes	yes	yes	yes
Household characteristics	yes	yes	yes	yes	yes	yes
Village characteristics and food prices	yes	yes	yes	yes	yes	yes
Province-specific year dummies	yes	yes	yes	yes	yes	yes
Number of observations	405	462	444	459	425	476

Source: China Health and Nutrition Survey, 1997–2006.
Notes: Standard errors in parenthesis are robust to heteroscedasticity and permit within-village correlations in unobservable.
**Significant at 5 percent.

6. CONCLUSIONS

In this paper, we study the effects of parent migration on young children's nutritional status in rural China. We find that migration is positively associated with short-term measures of nutritional status. According to our within-individual differenced and instrumented results, parent migration is associated with an increase of WAZ-scores of between 0.08 and 0.2 standard deviations. The estimated coefficients on HAZ and BMIZ-scores are positive but not significantly different from zero. These basic results are robust to different specifications and to a standard correction for panel attrition. In exploring possible mechanisms through which these impacts might occur, we provide suggestive evidence that migrant households are more likely to have tap water, which is likely to reflect an income effect.

Even though young children's nutritional status does not seem to be compromised by the absence of their parent(s), we are not optimistic that children left behind can benefit unambiguously from parent migration. First, our previous study (de Brauw and Mu 2011) shows that once children grow, they will have to take up more household chores. Any advantage they may have in WAZ when young will disappear. Second, the anthropometric measures, although easy to gauge, do not capture many important dimensions of health. For example, they cannot account for any psychological impact of parental migration on children.

However, it is important to note that migration seems to have a positive impact on household consumption of public goods, such as tap water. Further research is needed to analyze the mechanisms through which this could happen or more generally how migration is related to the provision and use of public goods.

APPENDIX: SUPPLEMENTARY TABLES

Table A.1—Summary statistics

	All	1997	2000	2004	2006
Children characteristics					
Age	4.62	2.75	4.42	6.14	5.27
	(2.38)	(1.57)	(2.13)	(2.29)	(1.81)
Boy	0.62	0.67	0.64	0.59	0.54
	(0.49)	(0.47)	(0.48)	(0.49)	(0.50)
Nutrition outcomes and inputs					
Height-for-age *z* score	–1.06	–1.12	–1.17	–0.86	–0.92
	(1.40)	(1.30)	(1.35)	(1.50)	(1.44)
Weight-for-age *z* score	–0.46	–0.45	–0.54	–0.44	–0.11
	(1.35)	(1.24)	(1.30)	(1.42)	(1.63)
BMI-for-age *z* score	0.18	0.41	0.17	0.02	0.24
	(1.41)	(1.36)	(1.37)	(1.43)	(1.66)
Calorie intake (kilocalories per day)	1240.49	1025.55	1272.56	1380.26	1075.3
	(561.91)	(431.43)	(568.68)	(605.73)	(439.80)
In-house flush toilet	0.18	0.14	0.14	0.21	0.28
	(0.38)	(0.35)	(0.35)	(0.41)	(0.45)
No excreta around house	0.55	0.48	0.54	0.59	0.66
	(0.50)	(0.50)	(0.50)	(0.49)	(0.48)
In-house or in-yard tap water	0.53	0.43	0.49	0.64	0.71
	(0.50)	(0.50)	(0.50)	(0.48)	(0.46)
Immunization last year	0.87	0.93	0.95	0.79	0.73
	(0.33)	(0.26)	(0.22)	(0.41)	(0.45)
Physical checkup past four weeks	0.03	0.04	0.03	0.03	0.02
	(0.17)	(0.19)	(0.17)	(0.18)	(0.15)
Household characteristics					
Head's years of schooling	6.78	6.7	7.02	6.66	6.11
	(3.39)	(3.52)	(3.28)	(3.32)	(3.79)
Number of females aged 60 and older	0.16	0.14	0.15	0.19	0.2
	(0.37)	(0.35)	(0.35)	(0.40)	(0.40)
Number of males aged 60 and older	0.12	0.08	0.1	0.17	0.18
	(0.33)	(0.26)	(0.30)	(0.37)	(0.38)
Number of working-age men	1.09	1.15	1.14	1.01	0.93
	(0.56)	(0.48)	(0.55)	(0.59)	(0.63)
Number of working-age women	1.2	1.27	1.23	1.14	1.09
	(0.56)	(0.61)	(0.52)	(0.55)	(0.59)
Assets per capita (log)	6.81	6.89	6.77	6.86	6.69
	(1.49)	(1.47)	(1.47)	(1.51)	(1.58)
Village characteristics and food price					
Free-market price of pork	10.13	11.35	9.15	11.29	7.78
	(3.15)	(3.21)	(3.20)	(2.28)	(2.31)
Free-market price of regular rice	1.92	2.33	1.65	2.04	1.83
	(1.73)	(2.86)	(0.99)	(1.64)	(0.54)
Free-market price of flour	2.05	2.16	1.79	2.45	1.63
	(1.92)	(1.65)	(0.60)	(3.16)	(0.35)
Village has a clinic	0.79	0.8	0.8	0.77	0.81
	(0.41)	(0.40)	(0.40)	(0.42)	(0.39)
Other households with tap water (%)	0.54	0.45	0.51	0.62	0.7
	(0.41)	(0.43)	(0.40)	(0.41)	(0.39)

Source: China Health and Nutrition Survey, 1997–2006.

Note: Standard deviations are in parentheses. Samples include children who were aged zero to five in 1997, 2000, or 2004 and were surveyed for two consecutive rounds.

Table A.2—Estimation of attrition probabilities: Children in the China Health and Nutrition Survey, 1997–2006

	Dependent Variable: Stay = 1, Move = 0		
	(2000|1997)	**(2004|2000)**	**(2006|2004)**
Height-for-age *z* score	–0.000	–0.014	–0.001
	(0.016)	(0.011)	(0.007)
Parent(s) migrated	0.019	–0.102	–0.048
	(0.124)	(0.098)	(0.054)
Child age	0.076	0.020	–0.024
	(0.065)	(0.063)	(0.043)
Boy	0.056***	0.041**	0.071***
	(0.020)	(0.019)	(0.015)
Head's years of schooling	0.006	–0.012	–0.004
	(0.010)	(0.010)	(0.007)
Number of females aged 60 and older	0.159*	0.003	0.062
	(0.082)	(0.118)	(0.076)
Number of males aged 60 and older	–0.205*	–0.118	–0.068
	(0.124)	(0.116)	(0.056)
Number of working-age men	–0.052	–0.173***	–0.050
	(0.062)	(0.060)	(0.040)
Number of working-age women	0.036	0.062	–0.003
	(0.061)	(0.071)	(0.032)
Household assets per capita (log)	0.025	–0.049*	–0.011
	(0.024)	(0.025)	(0.018)
Free-market price of pork	–0.014	–0.018	0.019***
	(0.010)	(0.020)	(0.006)
Free-market price of regular rice	–0.003	–0.023	–0.003
	(0.013)	(0.067)	(0.003)
Free-market price of flour	0.016	–0.001	0.045**
	(0.012)	(0.013)	(0.018)
Village has a clinic	–0.012	–0.014	–0.276***
	(0.106)	(0.082)	(0.095)
Other households with tap water (%)	–0.189*	–0.042	0.014
	(0.103)	(0.104)	(0.076)
Heilongjiang Province	0.176**	0.102	0.289*
	(0.086)	(0.117)	(0.169)
Jiangsu Province	–0.112	0.098	–0.133***
	(0.160)	(0.139)	(0.039)
Shandong Province	–0.228	–0.148	–0.028
	(0.174)	(0.169)	(0.096)
Henan Province	–0.245*	0.139	–0.086
	(0.134)	(0.117)	(0.071)
Hubei Province	–0.061	–0.149	0.060
	(0.118)	(0.123)	(0.122)
Hunan Province	0.044	–0.133	0.049
	(0.152)	(0.112)	(0.117)
Guangxi Zhuang Autonomous Region	0.014	0.038	–0.152***
	(0.110)	(0.115)	(0.040)
Liaoning Province		0.119	–0.155***
		(0.109)	(0.037)
	335	352	293

Source: China Health and Nutrition Survey, 1997–2006.
Notes: Marginal effects from probit model estimations are reported. Standard errors in parenthesis are robust to heteroscedasticity and permit within-village correlations in unobservable.
*Significant at 10 percent. **Significant at 5 percent. ***Significant at 1 percent.

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2033 K Street, NW
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Fax: +1-202-467-4439
Email: ifpri@cgiar.org