Migration and Young Child Nutrition:

Evidence from Rural China*

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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. In this study, we examine the impact of parental migration on the nutritional status of young children in rural areas. We use the interaction terms between wage growth in provincial capital cities and 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 it improves their weight. We provide suggestive evidence that the improvement in weight may be achieved through increased access to tap water in migrant households. Concerns about the sustainability of the impact on weight are raised in the conclusions.

Keywords: migration, children, nutrition, rural China

JEL codes: I1, J6, O1.

* We are grateful to Nicole Garcia and Feng Huang for great research assistance. For very useful comments we'd like to thank Joyce Chen, Kalena Cortes, Ranjita Misra, Hiroshi Ono and seminar participants at the 2012 American Social Sciences Annual Conference and the Asian Studies Forum, Texas A&M University.
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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 the impacts of emigration 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 households and communities they leave behind (Stark and Bloom, 1985). 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 et al., 2008; Yang, 2008). When migrants leave households, however, the time investment in raising children will also decrease (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). Therefore, migration can affect human capital outcomes of children in several different ways, and the overall effect depends on the relative importance of these factors.

A focus on children's nutrition is important because poor nutrition outcomes early in life have long term consequences for children (Grantham-McGregor et al., 2007). For example, multiple studies demonstrate that children's health and nutritional status have a sizable and statistically significant positive impact on later 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] Benefits are not limited to educational outcomes: longitudinal studies in developing countries also show evidence of improved cognitive measures later in life (Grantham-McGregor et al., 2007). Stunting before the age of 5 can also lead to less formal employment or lower wages later in life (Walker et al., 2011; Maluccio et al., 2009; Carba, Tan, and Adair, 2009).

Several recent studies suggest that migration can potentially influence nutrition among children left behind by the migrant or migrants. Gibson, McKenzie, and Stillman (2011) find that in Tonga, height-for-age z-scores (HAZ) are lower among children under 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,

[1] For a literature review of recent studies on the impact of child health and nutrition on education in less developed 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 her or his height in meters squared (weight [kg]/height$[m]^2$). BMI z-scores are used like weight-for-height z-scores but also control explicitly for age.

respectively, on child HAZ-scores, with the former finding positive impact but the latter a negative one. Differences in data, country contexts, child age definitions, empirical specifications, and methods may all 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). However, migrants are still largely barred, either legally or financially, from accessing education for their children in urban areas partially due to the household registration (*hukou*) system. 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 estimates suggest that as many as 23 million children (under age 14) are left behind in migrant-sending regions while their parents are working elsewhere (Guo, 2009).

This study builds on our previous analysis on 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 nonmigrant 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 (WHO, 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 very 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 on 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 on China's rural-to-urban migration and the nutritional status of 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 over the past three decades. In this section, we first outline the evolution of China's internal migration and its implications for rural communities, followed by a discussion on how rural children's nutritional status has changed over time.

2.1 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 one percent sample from the 1990 and 2000 rounds of the Population Census and the 1995 one percent population survey suggest that the intercounty migrant population grew from just over 20 million in 1990 to 45 million in 1995 and 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 over shorter periods of time (Cai et al., 2008).

The expansion of migration over 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 (e.g. Munshi, 2003). However, legal temporary resident status does not garantee 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 appear to be multifaceted. On one hand, remittances benefit migrant-sending households, as evidenced by their higher consumption (Du et al., 2006; de Brauw and Giles, 2008), improved risk-coping ability (Giles, 2006), and higher levels of investment in productive assets (Zhao, 2002; Woodruff and Zenteno, 2007). On the other hand, individual household members who are left behind, particularly women and seniors may be 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 et al., 2011). Even though children's educational performance is not compromised but 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 (Zhao, 1999; Rozelle et al., 1999). Since the late 1990s more women have participated in rural–urban migration, particularly the young and the single (Du et al., 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.

2.2. Nutritional Status of Rural Children

China faces multiple challenges in improving nutrition among rural children. First, disparities are present in child health within and among regions. For example, undernutrition 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 et al., 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 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. In 1990, the average height of children aged two to five in seven provinces was 92.5 cm 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 WHZ (weight-for-height z-score) increased dramatically between 1990

and 2000 (Chen, 2000; Osberg et al., 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 as unsafe drinking water can cause diseases such as diarrhea, which can severely affect child growth (e.g., Glewwe and Miguel, 2008). During the 1980s, the Chinese government launched a drinking water improvement program in rural areas. Clean water from water plants substantially improves rural children's HAZ and WHZ scores (Zhang, forthcoming). 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 migration improves household access to tap water.

3. Conceptual Framework

One can hypothesize at least three main channels through which migration might affect nutritional investment in children: the income effect, a time effect, and an effect from changes in the structure of intrahousehold bargaining and cooperation.

First, from the perspective of a household, the 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 to begin with (Fogel, 1994). The income elasticities of demand for 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). Diet improvments might manifest themselves in measures of nutritional status, particularly HAZ scores. 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 as well. 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, by sending out migrants, households decrease their overall labor endowment. Households in rural China are almost always engaged in some form of local income generation, either through farming or self-employment, therefore childcare time becomes scarcer in migrant-sending households. As a result, less time is allocated to cooking (Chen, 2006; de Brauw and Mu, 2011) or monitoring the eating habits of children. To the extent that it is difficult and time-consuming to ensure nutritional

[4] To be precise, we consider the migrant still a household member after migrating, at least until she or he 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 his or her past income contribution.

investments in children, the time effect of migration may compromise child nutrition, especially among younger children.

Third, when household members live apart as a consequence of migration, the equilibrium in the implicit intrahousehold 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 (e.g., 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 innocuous impact on child nutrition (Chen, 2006).

As outlined in this conceptual framework and suggested by mixed results from existing studies, whether or not 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 measure this relationship in rural China.

4. Data and Empirical Framework

4.1. Data

The data used in the analysis are primarily derived from four waves of the China Health and Nutrition Survey (CHNS), collected in 1997, 2000, 2004 and 2006.[5] The CHNS is a longitudinal survey covering nine provinces that vary substantially in

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

geography, economic development, and access to public resources.[6] Among other information, the survey includes the demographic characteristics and asset holdings of each household. We further make use of community-level data to account for various village charateristics.

Important for this paper, the CHNS records: height and weight for each individual within the household, measured by a physician, nurse or a health worker during the survey. To construct HAZ, WAZ, and BMIZ scores, we use the most recent growth charts made available by the WHO, which have corrected previous standard growth charts for what are now considered developed-country biases (de Onis et al., 2007).[7]

The migration status of household members in CHNS can be built up from the household roster. Starting from 1997, if an individual who was in a previous round of the CHNS is not residing in the same household in the current survey, a question is asked regarding the reasons for his/her absence. We consider any individual who has left the home to seek employment at the time of the survey to be a migrant. This definition differentiates migration for work from migration for other reasons, but it could undereport the migration rate because return migrants cannot be identified. Hence, we may underestimate the impact of parents' migration on children's nutrition status. On average, at the time of each survey migrants had been away from home for about 14 months. As our instrumental variables are suited for identifying the impact of migration per se and not of the lengths of the migration spell, our results below provide estimates of the average impact of migration.

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

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

We pool panel data from all three pairs of surveys—1997 and 2000; 2000 and 2004; 2004 and 2006—and keep sample children aged under 5 years in the initial years (1997, 2000, and 2004) and surveyed 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, 1995). 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 (Figure 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, consistent with findings in other studies (e.g., 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 above 20 percent in 2006. It could reasonably be expected that migration by individuals in these cohorts have the largest chance of affecting the nutritional status of children.

Consistent with the evidence previously presented in Figure 1, over time more children in our sample are found to be living in migrant households (Table 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 in 1997 to 13.3 percent of households 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 a sole migrant.

In Figure 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, the average HAZ is under -1, implying a significant stunting rate.[8] By 2006, the average HAZ score increases to around -0.7, with the scores of the very young children approaching those of the healthy reference population. The age profile of HAZ shows that this measure deteriorates from birth to approximately 48 months. After that the average HAZ levels off or improves slightly. This pattern is similar to that found in Guatemala (Carletto et al., 2011), where the turning point is around 30 months. WAZ is negatively correlated with age, a pattern common in many developing countries, such as El Salvador (de Brauw, 2011). A steady increase in WAZ is evident during the period between 1997 and 2006. Children younger than 3 have already had average WAZ equal to or above 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.

[8] Children with an HAZ below -2 are considered "stunted."

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 3 shows that the middle of the distribution of HAZ for children whose parents did not migrate is positioned slightly right to that for children with a migrant parent, but no distinct shift in the whole distribution is observed. The distributions of WAZ and BMIZ exhibit a similar pattern for these two groups of children.

4.2. 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 H_{ihjt}. It is a function of parents' migration status (M_{hjt}), which is measured as 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 in the initial period (A_{ihj0}); household (X_{hjt}) and village (V_{jt}) characteristics; province-specific year effects $Y_{p\times t}$; and individual fixed effects (μ_i). The inclusion of $\mu_i$effectively controls for such important determinants of child height and weight as maternal height (Ozaltin, Hill and Subramanian, 2011). To account for the way children grow as they age in the early years, we allow age and gender effects to vary over time by interacting A_{ihj0} with a time trend T. With the idiosyncratic error term denoted as ε_{ihjt}, the estimation equation can be written as follows:

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

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 very 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 above) 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 as through relative wealth and household diet, food prices may influence migration decisions as well as 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 panel data, 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 must still be concerned that unobservables that vary over time may be important determinants of child nutrition as well as 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 causal impact of migration, we must find variables that affect migration but do not independently affect children's nutritional status. In this context, we suggest the following identification strategy. We hypothesize that wage growth in provincial capital cities (ΔW_{pt}) potentially affects household migration decisions. A large share of migration during 1990-2000 is intraprovincial (72.3% in 1990 and 46.1% in 2000) (Liang and Ma, 2004). Even though the interprovincial migration, responsive to the increasing interregional disparities, has gone up over time, according to CHNS more than one third of the migrants move within provincial borders. Wage growth in the provincial capital city therefore could proxy for the demand side signal that migrants would observe. The CHNS community survey contains information for each primary survey unit on the daily wage for an ordinary male and female worker separately. Using the mean reported wages for each capital city, we calculate the wage growth between survey years.

We further hypothesize that the strength of the wage signal in citites will also be affected by the initial size of the household migrant network (M_{-hj0}), as measured in the 1997 survey.[9] More specifically, the migrant network is measured as the share of men and women in the working-age population who had migrated in 1997. We exclude the household when calculating its network. The change in migration can therefore be written as follows:

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

[9] The reasons for absence were recorded only from the 1997 round onward, 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.

It is conceivable that wage growth and the migrant network individually affect migration decisions; each variable on their own might also be correlated with children's nutrition outcomes even after we account for migration. For example, wage growth in the provincial capital might be correlated with changes in local wage , which could be reflected in changes in anthropometric measures through the income effect. Households with large migrant networks might have better access to nutrition information affecting anthropometric outcomes independently of migration. Once we control for the migration network and the provincial capital wage growth directly in the regression, the variation in the interaction term is unlikely to independently affect children's anthropometric outcomes. In equation (2), we implicitly control for wage growth (ΔW_{pt}) through the province-specific year effect ($\Delta Y_{p\times t}$) and migration network (M_{-hj0}) through fixed effects. The interaction term alone ($\Delta W_{pt} \times M_{-hj0}$) is therefore used for statistically identifying migration.

Given wage growth, migration is more likely among individuals with larger migrant networks, and their children are more likely to be left behind. This relationship is confirmed in Figure 4, in which we stratify villages into two groups: a group with larger migrant networks and a group with smaller migrant networks. We define the cutoff for large migrant networks as villages with at least 15 percent of working-age women or 20 percent of working-age men migrating in 1997. The graphs illustrate that in villages with larger migrant networks, migration responds positively with wage growth, but the relationship between migration and urban wage growth is much weaker in villages with smaller migrant networks. Moreover, this pattern holds for both men and women. This approach essentially exploits the complementarity 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.

5. Results

5.1 Primary Results

We first examine the HAZ score as the dependent variable and estimate equations (1) and (2), respectively, treating parental migration as exogenous. The results are reported in Table 2. In the OLS regression, the 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 statistically significant when we take the first difference. Two estimations yield small coefficients with large standard errors on the indicator variable for parental migration. 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.

Next, we use the interaction terms described above as instrument variables for migration, and we reestimate equation (2) using an instrumental variables estimator (Table 3). 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 overidentification test is passed comfortably, but under standard assumptions about the error terms, the instruments would be considered weak (e.g. Stock and Yogo, 2005). In the presence of clustering, it is less clear though whether these F statistics suggest weak instruments. We therefore provide the 95 percent confidence interval based on Moreira's (2003) weak instrument robust conditional likelihood ratio statistic. The statistic is adjusted for clustering and robust to weak instruments (Finlay and Magnusson, 2009). The interval for the HAZ result suggests that the null of zero impact cannot be rejected at the 95 percent significance level.

In the second stage, we again find no significant relationship between the migration variable and child HAZ scores (Table 3, columns 2). As expected, the standard errors rise, whereas the point estimate remains small. Consistent with results reported in Table 2, these results confirm that parental migration does not affect children's HAZ scores.

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 4. 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 a statistically significant coefficient of reasonable magnitude: 0.201 standard deviations. This estimate

essentially suggests that the children of migrants are heavier than children of others, *ceteris paribus*, at the same age among children of the same gender. Even if the point estimate suffers from weak instrument bias, the 95 percent interval is positive throughout its range, suggesting migration has a unambiguously positive impact on WAZ scores. Given the absolute lack of significance when using HAZ scores as the dependent variable, migration does not appear to affect the long-term nutritional measure, but the short-term measures. In fact, the 95 percent confidence interval for the coefficient estimate on BMIZ is primarily positive, adding support to this argument. Therefore, our inability to reject the null hypothesis that the coefficient estimate of the BMIZ scores is zero is likely the result of a lack of statistical power in the regression.

In summary, our primary regression results indicate that parental migration has a positive effect on children's weight but has no impact not 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.

5.2. 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.

Compared to the previous estimates with household level control variables, the coefficients on migration in Table 5 are smaller in all three regressions. One possible reason for this 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 nutritional outcomes.[10] 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.

We are also concerned about potential panel attrition bias. The CHNS sample attrition is high for children aged 0 to 5, particularly between 2004 and 2006 (Table 6). 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 6 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. With respect to these observables the panel attrition would therefore appear to be random.

Nonetheless, we reestimated 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 (results reported in Appendix Table 2). We subsequently use the inverse of that estimated probability as a weight in the

[10] The number of female household members aged 60 and above and the number of working-age women have positive and significant coefficients in the BMIZ and WAZ regressions. These results are not reported in Table 4; they are available upon request.

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 (Table 7), we find that the results are largely unchanged. We find statistically significant estimate for the effect of migration on WAZ scores but not BMIZ or HAZ scores.[11]

5.3. 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 (Table 8).[12]

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 8). 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

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

[12] 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.

percentage points more likely to have a water tap in the house or in the yard by the second survey relative to nonmigrants. 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.

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.

The estimates are average impacts of migration. Heterogeneities in terms of the timing of migration are not studied in this paper, but deserve further research. Even though our results indicate that 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 may 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.

Figure 1. Migration rate by gender and year: China Health and Nutrition Survey, 1997-2006

Figure 2. Nutritional outcomes of children aged 0-9 by year: China Health and Nutrition Survey, 1997-2006

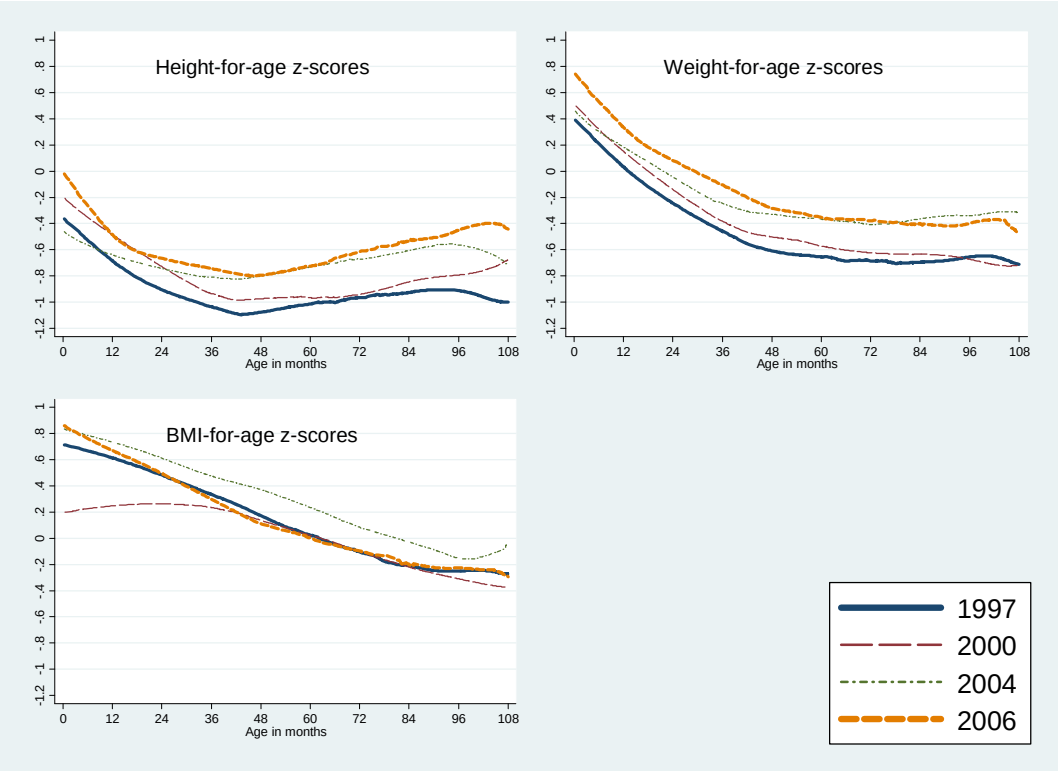


Figure 3. Nutritional outcomes of children Aged 0-9 by parent migration status: China Health and Nutrition Survey, 1997-2006

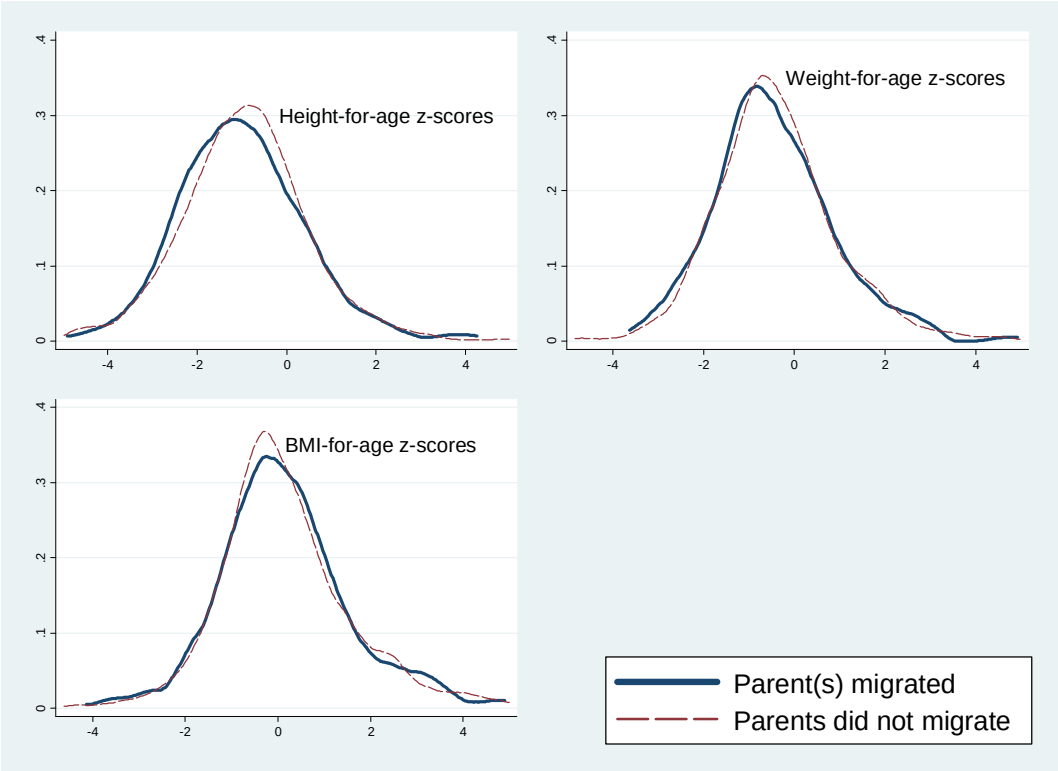
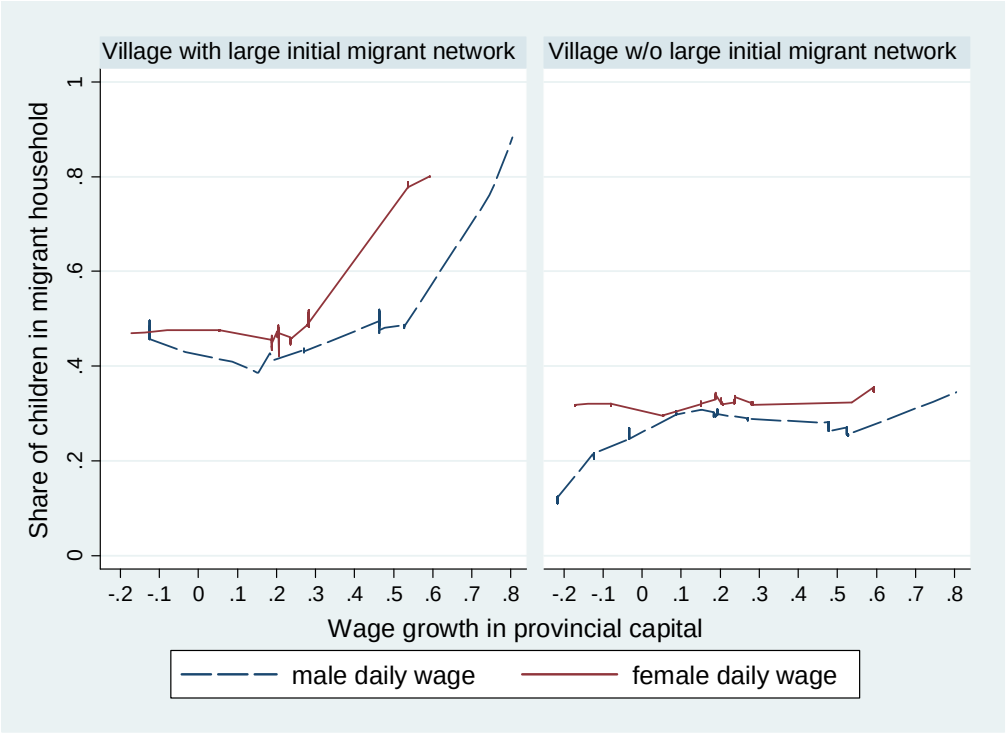


Figure 4. Share of children in migrant households and wage growth in provincial capital, by village initial migrant network: China Health and Nutrition Survey, 1997-2006



Note: "Villages with large initial migrant network" are defined as villages in which 15% of working-age women or 20% working-age men migrated in initial years. The rest of the villages are labeled as "villages without large initial migrant network".

Table 1. Migration of parents of rural children: China Health and Nutrition Survey, 1997-2006

	1997 (aged 0-5)	2000 (age 0-8)	2004 (age 0-9)	2006 (age 2-7)
Any Parent Migrated	0.060	0.116	0.231	0.337
	(0.238)	(0.320)	(0.422)	(0.476)
Father Migrated Only	0.034	0.069	0.090	0.133
	(0.181)	(0.254)	(0.286)	(0.341)
Mother Migrated Only	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 Obs.	239	479	325	85

Note: Standard deviations are in parenthesis. Rural children who were aged 0-5 in 1997, 2000 or 2004 and surveyed in CHNS for at least two consecutive rounds with non-missing anthropometric measures are included in the calculation.

Table 2. Baseline regression results on children's height-for-age z-score, China Health and Nutrition Survey, 1997-2006

Dependent variable: HAZ	OLS	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)
No. of female aged 60 and above	0.335**	0.058
	(0.153)	(0.186)
No. of male aged 60 and above	-0.012	0.227
	(0.191)	(0.239)
No. of working-age men	-0.081	0.098
	(0.103)	(0.105)
No. 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

Notes: Standard errors in parenthesis are robust to heteroskedasticity and permit within village correlations in unobservables. *** significant at 1%; ** significant at 5%; * significant at 10%.

Table 3. Instrumental variables estimates of the determinants of children's HAZ, China Health and Nutrition Survey, 1997-2006

Dependent variable: ΔHAZ	**First Stage**	**Second Stage**
ΔParent(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	
Prob > F	0.001	
Over identification: Hansen J statistic	0.109	
Chi-sq(1) p-val	0.741	
Confidence intervals derived from weak-instrument robust tests	[-0.090, 0.284]	
Number of observations	491	

Notes: Standard errors in parenthesis are robust to heteroskedasticity and permit within village correlations in unobservables. *** significant at 1%; ** significant at 5%; * significant at 10%.

Table 4. Instrumental variables estimates of the determinants of children's BMI-for-age and weight-for-age z-scores, China Health and Nutrition Survey, 1997-2006

	ΔBMIZ	**Δ WAZ**
ΔParent(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
Prob > F	0.001	0.000
Over identification: Hansen J statistic	1.406	0.306
Chi-sq(1) p-val	0.236	0.580
Confidence intervals derived from weak-instrument robust tests	[-0.008, 0.301]	[0.019, 0.520]
Number of observations	455	511

Notes: Standard errors in parenthesis are robust to heteroskedasticity and permit within village correlations in unobservables. *** significant at 1%; ** significant at 5%; * significant at 10%.

Table 5. Robustness check: instrumental variables estimations without household characteristics, China Health and Nutrition Survey, 1997-2006

	ΔHAZ	ΔBMIZ	Δ WAZ
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
Prob > F	0.000	0.001	0.000
Over identification: Hansen J statistic	0.057	1.406	0.306
Chi-sq(1) p-val	0.812	0.236	0.580
Confidence intervals derived from weak-instrument robust tests	[-1.463, 0.868]	[-0.368, 0.408]	[0.015, 0.224]
Number of observations	491	455	511

Notes: Standard errors in parenthesis are robust to heteroskedasticity and permit within village correlations in unobservables. *** significant at 1%; ** significant at 5%; * significant at 10%.

Table 6: Anthropometric measures among children and parent migration status by panel attrition, China Health and Nutrition Survey, 1997-2006

	Mean Difference by Attrition		
	Stayers	**Leavers**	**(1)-(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	

Note: Sample means are reported in column (1) and (2). The last column reports t-statistics for testing the equality of the two sample means.

Table 7. Robustness check: instrumental variables estimations with inverse probability weighting to correct for panel attrition, China Health and Nutrition Survey, 1997-2006

	Δ HAZ	**ΔBMIZ**	**Δ WAZ**
ΔParent(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
Prob > F	0.003	0.000	0.000
Over identification: Hansen J statistic	0.001	2.514	0.303
Chi-sq(1) p-val	0.973	0.113	0.582
Confidence intervals derived from weak-instrument robust tests	[-2.543, 1.256]	[-0.008, 0.284]	[0.019, 0.520]
Number of observations	491	455	511

Notes: Standard errors in parenthesis are robust to heteroskedasticity and permit within village correlations in unobservables. *** significant at 1%; ** significant at 5%; * significant at 10%.

Table 8. Instrument variables estimates of impacts of migration on hygiene environment and health inputs: China Health and Nutrition Survey, 1997-2006

	Δ Calorie intake (Kcal per day (log))	Δ In-house flush toilet	Δ No excreta around house	Δ In-house or in-yard tap water	ΔImmu-nization last year	Δ Physical check-up past 4 weeks
Δ Parent(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
F test on excluded instruments	8.38	5.52	5.19	5.54	7.81	8.11
Prob > F	0.0004	0.005	0.006	0.005	0.001	0.001
Over identification: Hansen J statistic	4.165	0.661	7.044	0.018	0.023	0.000
Chi-sq(1) p-val	0.041	0.416	0.008	0.894	0.881	0.986
Confidence intervals derived from weak-instrument robust tests	[-0.950, 0.376]	[-0.116, 1.497]	[-0.310, 1.226]	[-0.004, 0.117]	[-0.921, 0.509]	[-0.390, 0.066]
Number of observations	405	462	444	459	425	476

Notes: Standard errors in parenthesis are robust to heteroskedasticity and permit within village correlations in unobservables. *** significant at 1%; ** significant at 5%; * significant at 10%.

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