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Child Nutrition and Vietnam's Transition to the Market

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ABSTRACT: This paper measures the social cost of economic transition in terms of child health, as measured by nutritional status. It sets to answer the question if children are better-off or worse-off nutritionally during the early stages of market openness. Focusing on Viet Nam, we explore this question empirically by developing a pseudo panel approach using household cross-sectional living standards data. We find that in the short-run, market openness did not improve child growth.

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JEL classification: I12; P36

Key words: market transition; child health and nutrition; household behavior

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1. INTRODUCTION

Recent years have witnessed the transition to a market economy in former centrally planned economies. These transitions have been both dramatic--the collapse of Eastern Europe and the former Soviet Union, and evolutionary-- Communist China's gradual direction to a market economy (Lipton and Sachs 1991, Kornai 1992, Rühl and Serwin 1994). In various degrees, both transitional approaches to market openness have had profound effects on the welfare of households (Feachem 1994; World Bank 1996). Market openness promises a better life through increased choice in jobs, goods and services, education and health care through flourishing market opportunities and competition. But for some households, market openness has had painful consequences--increased risk and uncertainty in employment, real incomes and savings eroded by inflation and declining social protection (Hiebert 1993; Leijonhufvud 1993; Hiebert 1994; Sipos 1994). As a result, in most early transformers, despite economic growth, poverty has risen, the gaps in income between regions and households have widened, and the health of vulnerable groups, particularly children, are at risk (Cornia and Sipos 1991; Shen, Habicht and Chang 1996). Governments embracing market openness must consider the impact of market transition on the welfare of their citizens. Yet at these early stages of reform, governments have very little information to track household welfare over time.

This paper is about the measurement of the social cost of economic transition in terms of child health, as measured by nutritional status. It sets to answer the question if children are better off or worse off nutritionally during the early stages of market openness. Focusing on Viet Nam, we explore the impact of *doi moi*, (a regime change toward market openness in December 1986), on child growth.[1] Unlike most transformers, who experienced an early lag period in economic growth, Viet Nam's economy responded quickly to the *doi moi* openness regime and grew an average of 8% per year in the 1990s. Given this economic growth and the fact that the initial reforms liberalized the food market, and increased farm incomes for this largely agricultural nation, we hypothesize that market openness in Viet Nam had a positive effect on child growth. We explore this question empirically by developing a pseudo panel approach and applying this approach on the World Bank's Viet Nam Living Standards data 1992-93, a cross-sectional dataset consisting of 4800 households.

The paper begins with a description of the institutional setting in Viet Nam before and after *doi moi*. The following sections discusses the previous work in the nutrition-income literature in developing countries, the data, the methods, and estimation issues. The final sections present the model results, discuss the interpretation of the findings within the country context, note the limitations of the estimated model, and propose inferences that can be derived from the results for public policy.

2. INSTITUTIONAL SETTING

"Market openness" in Viet Nam symbolically began with the socialist republic government's announcement in December 1986 to initiate a *doi moi* (economic renovation) program, although radical economic change did not begin till 1987. The country embraced a regime typified by classic market openness reforms (Lipton and Sachs 1991) principally by liberalizing price and volume controls, adhering to hard budget constraints and privatizing commune-owned enterprises. Over the last decade business pundits have been bullish about Viet Nam—boasting economic growth rates averaging 8% per year in the early 1990s (World Bank 1993b).

Viet Nam's growing economy seems to have had an overall positive effect on reducing child malnutrition. During the period 1982-1985, with malnutrition approaching 60%, Viet Nam's economic

[1] Height-for –age Z score.

growth had stagnated, worsened by economic sanctions from China, the ASEAN countries, the United States, and other Western countries (World Bank 1993a). Malnutrition rates improved by 4% in 1987-1989, marking the beginning of market reforms and the lifting of trade embargoes (World Bank 1995a). By 1992-93, Viet Nam's economy was growing rapidly and the malnutrition rate dropped from 56% to 52.5%.[2] (World Bank 1995b). Despite this decreasing trend, malnutrition in Viet Nam is still one of the highest rates world wide. Viet Nam's malnutrition situation among children can be summarized as a vicious cycle of poor food security for most families, inadequate maternal intakes that lead to a prevalence of low birthweight babies, inappropriate breast feeding and weaning practices and an insufficient and unbalanced diet (Haughton and Haughton 1997).

The post-*doi moi* period encapsulated a regime change of several polices affecting internal and external markets in labor, trade, and money. The regime change also included devolving social protection in education and health to local private markets. Thus, many aspects of *doi moi* had consequences on household resources, and hence the nutritional status of the child. However, we posit that two fundamental reforms had the most direct consequences on child health. First, decollectivization of communal farms in 1988 opened up food markets and presented the opportunity for farm households to increase income (Fforde and de Vylder 1996, Hiebert 1993b). Second, in 1989, the government opened the door to private markets in health hoping to mobilize private resources and introduce more choice for consumers through provider competition (Gellert 1995; Hein et al. 1995).

Viet Nam's recent emergence as a major rice exporter is credited to the immediate productivity gains following the decollectivization of agricultural land. With the reforms, farm households had a direct incentive to maximize their plot yields. Decollectivization gave immediate food security benefits to farm households, with implicit benefits to household nutritional status. (Jamison, Nguyen and Rambo 1992)

While nutritional gains are expected from the opening of food markets, it is less clear if child growth would benefit from a more open health care market. In March 1989, the government legalized private practice for physicians and physician assistants, enabling them to either set up full private practice, or conduct off-hours private practice in community health facilities. A month after, the government liberalized price and output decisions in the pharmaceutical market. By May of 1989, user fees were introduced for district, provincial and national health services and in June 1989, the government legalized retail pharmacies. (Shellard 1992; World Bank 1995b).

Consumers of health care, particularly those who had lived their entire lives under the socialist net of health care, had little clue of the fair price to pay. Consequently, some physicians and pharmacists were charging steep, usurious fees to patients constrained by transportation costs and the need for urgent care. Because there is currently no systematic regulation of prices for physician services and pharmaceuticals, consumers could face great price differentials by province and by urban versus rural (Nguyen et al. 1995; Truong et al. 1995)

It is not clear what the impact of the restructured health sector has been on the nutrition situation. Competition has probably increased production efficiencies and consumer choice. User fees have mobilized resources to maintain the upkeep of deteriorating public health facilities. But currently, the unregulated public-private mix seems to have produced wage and price disparities in the health care market. Moreover, because 75% of total health expenditures are now privately incurred (compared to 0% pre-*doi moi*), poorer Vietnamese may be priced out of the system. Indeed, utilization rates for the poorest quintile have decreased over the past five years. (Gertler and Litvack1998). Decreased access to prenatal care, health education, and growth monitoring and immunization programs could compromise the nutritional status of the child, and pregnant and lactating mothers. Moreover, since pharmaceutical

[2] The prevalence rates of malnutrition are for children under 5 years old.

expenses are now 100% out-of-pocket, poorer families may be less likely to purchase necessary but expensive prescriptions that palliate illnesses and infections that slow child growth.

To summarize this section, there are aspects of Viet Nam's post-*doi moi* economy that have both bullish and bearish consequences on child growth. On the one hand *doi moi* jump-started the economy into high growth rates and presumably higher household incomes. Higher incomes have led to improvements in food security of the household and better nutritional status of its members. Moreover, freeing up price controls for food greatly contributed to economy-wide growth and probably increased domestic consumer choice and access to food. On the other hand, there is evidence that suggests that opening up health care markets, particularly at the commune level, may have been detrimental to child growth, particularly for the poor. Two issues arise. First, because of the possible offsetting effects of market openness in the food market and the health care market, it is less clear that the net effect of market openness (*doi moi*) will be remarkably positive. Second, since the economy was growing in Viet Nam during economic transition (and not so for most other reformers), the effect of market openness may be overstated. Hence, the income effect needs to be disentangled from the market openness effect on nutritional status. To shed light on this issue, the next section reviews previous work in the nutrition-income literature.

3. PREVIOUS WORK

While transition economies is a recent phenomena, this study's investigation on the effect of markets on nutritional well-being of children fits in the genre of work that looks at poverty and malnutrition, or more broadly, economic development and nutritional status.

Recent studies have borrowed from the earlier work on child survival (Mosley and Chen 1984) to develop frameworks for determining child health outcomes (e.g. diarrhea and anthropometry). Mosley and Chen's biological pathways, coupled with Becker's (1991) treatise on the economics of the household and Grossman's (1972) concept of health as both an investment and consumption good, led to studies which integrated biological and behavioral frameworks governed by household decisions (Da Vanzo and Gertler 1990, Thomas and Strauss 1992, Barrera 1990, Cebu Study Team 1992). These studies have used empirical data to estimate health production functions that encompass both physiological processes and the underlying environmental and economic conditions that influence child health. There are two advantages to this genre of work: (1) unlike most biomedical models, they typically control for the endogeneity of causal factors; and (2) unlike most economic models, they attempt to identify the biological determinants of health status. However, the disadvantage of these studies is that most models are empty of indicators that could evaluate explicit policy changes or public health interventions. Some studies convey limited understanding of the health outcome measure, resulting in misspecification of relationships and errors in measuring identifying variables. Despite these disadvantages, these studies are rigorous, a product of refinements in empirical work over the past fifteen years.

Anthropometry, as a measure of health outcome, represents objective measures of health and nutritional status (WHO 1995; Waterlow 1994). Height-for-age versus weight-for-age is a better stock measure of health because it measures a child's cumulative nutritional status since birth (Healy 1986). Moreover, because a chief concern in child nutrition is the failure to reach the growth potential as a result of inadequate nutrition or poor health, the indices of height-for-age are usually compared with a reference population of healthy children.[3]

[3] The World Health Organization (WHO) recommends the use of a United States reference sample collected by the National Center for Health Statistics (NCHS).[3] Normative measures, particularly those imposing industrialized country standards on the rest of the world have usually ignited decades-long debates. However, the controversy in imposing United States reference standards (NCHS/WHO) to evaluate children's growth has been put to rest by recent work (Martorell, Mendoza and Castillo 1988 and Martorell et al. 1995) in favor of the use of international standards. Martorell et al.'s 1995 work documents that similar growth patterns are observed among top quintile

Thomas, Strauss, and Henriques (1990) use height of children under eight years (as percentage of the median of standard growth tables for the child's age and sex) for the outcome variable and control for background variables by including mother's height and father's height to control the effect of mother's education. Thomas, Strauss, and Henriques (1990) instrumented per capita consumption with non-labor income to purge its endogeneity with nutritional status, but did not use a fixed effects estimator, in part because they stratified the Brazilian sample into four regions. Thomas and Strauss (1992) augmented their work by merging community characteristics with the Brazilian household sample to test their hypothesis on the effect of community infrastructure on child height. Unlike the earlier Thomas, Strauss, and Henriques (1990) work, income was modeled as per capita expenditure, versus total expenditures, conceding to inclusion of family size despite its endogeneity. Dividing total expenditures by family size allows for more precise estimates of individual allocation of resources within families, though fertility and income decisions are deemed to be jointly determined (Horton 1986). In later work, the virtue of including family size has prevailed. Like Thomas' previous Brazil study, mother's and father's heights were included as they are exogenous, predict child height well and control for background characteristics that may be overstated in the parental education variable. However, selection bias may be overlooked as only families that did not include both parental heights were excluded from the sample.

While family background, suitably controlled by the latest genre of work, is also important, idiosyncratic differences of children, like birth order, have tended to be overlooked by past studies (Horton 1988). Resources may flow more to the oldest child than to the youngest child. Also, exclusion of single-parent families may exclude families that are systematically different from families with both parents. Families without fathers may earn less income and have fewer time resources available for raising the child. This would be pronounced in Viet Nam, as there are a sizable number of female-headed households (Hiebert 1994b).

Studies using child anthropometrics could be improved with more objective measures of nutritional status. Percentage of standardized median, used by Thomas, Strauss, and Henriques (1990) and Thomas and Strauss (1992) was a step toward this direction but, their choice of measure fails to capture the normative assessment of malnutrition across different age groups.. Percentage of median, a choice measure for population studies a decade ago, is the expression of the observed child's height (or weight) as a percentage of the reference group's median height (or weight) for a given age and sex. While this allows for reporting of means and standard deviations, (i.e. accounts for the distribution of the sample) the nutritional status inference is inconsistent for different age groups (WHO 1995). To illustrate, we compare an infant who is 80% of the median for her age group with her older brother who is 75% of the median for his age group. From this statistic, it may seem that the infant is better off than her brother, but in truth, the infant may be 2 standard deviations below the median, therefore malnourished and her brother only 1 standard deviation below his reference median. Since this study compares the nutritional stock of siblings spanning the ages of 0 to 10 years of age, Z-score was the compulsory index of choice.[4] In a recent study on the effect of market reforms on child nutritional status on China, the authors use HAZ score as the outcome measure (Shen, Habicht and Chang, 1996).

Guatemalan children and Mexican-American children. The consensus, to date, is for pre-pubescent children (i.e. children under 10), population variations are largely due to environmental, rather than genetic differences (WHO 1983).

4 By using reference standards, height-for-age can be expressed in terms of Z-scores, which indicate, in terms of standard deviations, how far a child's height-for-age is from that of an average healthy child of the same age (or same height). A Z-score of -2 or below signals malnutrition; in a population of healthy children only 2-3 percent would have such a low Z-score. A Z-score of -3 or below indicates severe malnutrition, since less than 0.5 percent of children in a healthy population would have such a low Z-score.

Though the nutrition-income literature improves the estimates of income nutrition elasticities and in understanding the technology used at the household level, addressing broad policy questions such as evaluating interventions and structural changes has been elusive. This is in part due to the analytical restrictions of cross-sectional data used by most of these studies. This work will advance the empirical work through better identification and specification of the household production of child health. Further, by introducing a pseudo panel approach with cross-sectional data, we advance the static limitation of the nutrition-income literature by evaluating policy-induced changes in the nutrition situation over time.

4. DATA

The Viet Nam Living Standards survey (VNLSS) was conducted from October 1992 to October 1993. The nationally representative, self-weighted sample was drawn using a 3-stage sampling procedure: stage 1) selection of 150 communes out of Viet Nam's 10,000 communes, stage 2) selection of 2 villages within each sampled commune and stage 3) selection of 16 households from each of the 300 sampled villages. The overall survey sample consists of 4800 households comprising 23,839 individuals. (World Bank 1995a). We restricted the empirical analysis to children under age ten because these relatively young children display limited genetic variation during this pre-pubescent period (Martorell et al. 1995). The sample contains 4800 households with a total of 5652 children. About 84.5% of the sample are rural children.

For this analysis, nutritional status is specified as height-for-age Z score (HAZ), computed as follows (Gibson 1990):

$$HAZ = \frac{(H_c - H_r)}{SD_r} \tag{1}$$

where H_c is the height of a child, H_r is the median height of the healthy reference child of the same age and SD_r is the standard deviation of the distribution of heights for the NCHS/WHO reference population of healthy children of that age. Outliers with HAZ scores greater than |-7| were excluded from the sample. These were most likely due to errors in field measurement. Outliers constituted 14 out of 5652 cases. An additional 78 observations were excluded due to missing observation. The final sample for analysis contains 5560 observations.

Table 1 presents summary statistics of relevant variables in the dataset. Table 2 displays the mean HAZ-scores by groups of characteristics. HAZ scores lower than -2.00 signal moderate to severe stunting and that HAZ scores between -1.5 to -2.0 signal mild to moderate stunting. Among the regions, the Southeast has the best average Z-score (-1.55) and the North Central has the worst (-2.34). Urban children are better off than rural. Higher parental educational attainment seems to not have any marked effects on improving nutritional status of their children. Among ethnic groups, Chinese, on the average, have the best HAZ score. Females are slightly better off than males. Oldest children have better mean nutritional status than non-oldest children.

5. METHODS

This section discusses the methods used to test the hypothesis that market openness improved child growth in Viet Nam. Since household incomes were also growing after *doi moi*, the estimation strategy is to separate out the effect of income from market openness and to estimate an equation that minimizes unobserved heterogeneity within families. We begin with the specification of the model derived from the household production of child health, extend this model by identification of a market openness variable, and translate the identification discussion into the development of a pseudo-panel model.

5.1 Model Specification

Following Grossman (1972), health can be thought of as a stock of human capital. An individual's stock at a point in time is determined by an initial genetic endowment, subsequent behavioral choices (such as levels of work activity and medical care choices) and effects of the public health environment, such as living in an area where malaria is endemic (Cebu Study Team 1992). Over a period of time, the change in a person's health status is determined through a production function that transforms inputs into health. Some of the inputs are chosen. These include market goods (e.g. food and antibiotics) and time (food purchase and preparation time, time for medical consultation). Other inputs are not chosen by the household, such as the portion of the disease environment that is determined by public health and sanitation infrastructure.

The productivity of these inputs depends not only on biological mechanisms, but also on individual and household characteristics such as age, gender, education and family background (Da Vanzo and Gertler 1990; Thomas and Strauss 1992). For example, better educated individuals are better able to follow medical treatment procedures and consequently their use of medical care is more productive. Also, community factors, such as the quality of medical care, affect the productivity of the inputs. Thus, households make decisions so as to maximize their overall welfare, conditional on their resources.

Equation [1] is the health production function subject to a household budget constraint (income) I_f. ***H*** represents the individual child's health, where the child's nutritional status is one dimension of assessing the healthiness of the child. The child's current health is determined by: her initial health stock, ***Ho****;* a feasible combination of technological inputs, ***B,*** which could be biological (immune response to diseases, efficiency of nutrient absorption, rate of metabolism), or behavioral (was the child immunized, is her growth being monitored, was the child breast-fed, was the child weaned properly); exogenously determined inputs, $\boldsymbol{E_i}$, the characteristics of the child (age, sex), E_f, the characteristics of the family (mother's and father's education, age of mother at childbirth, parental height) and E_c the characteristics of the community (commune, region, urban/rural, time to nearest health services, average prices of pharmaceuticals); and all other unobserved factors, ϕ.

$$H = H(H_0, I_f, B \ , E_i, E_f, E_c, \phi) \qquad (1)$$

Though accounted for in the health production function, biological and behavioral inputs, ***B***, are difficult to observe and measure. Following Thomas and Strauss 1992, Equation [1] can be simplified into the reduced form of the health production function. The reduced form has a key advantage over the structural form, Equation [1]. The reduced form bypasses the estimation issues arising from the recursive processes of biological and behavioral adjustment. However, there is a serious disadvantage with the reduced form. The reduced form loses behavioral information and disease environment risks that could guide health policy makers in designing appropriate interventions, such as health education and anti-malaria campaigns. Yet for the scope of this research, the reduced form is both necessary and sufficient in that the aim is to estimate a consistent and unbiased estimate of income and market openness without the confounding endogeneity effect of biological and behavioral mechanisms, ***B***. The reduced form model of the child health production function is:

$$H_i = \mathrm{H}(H_{i0}, I_f, E_i, E_f, E_c, \theta) \qquad (2)$$

where H is child health, i indexes the individual child**,** f indexes the family/household, and c indexes the community. Nutritional status is one dimension of child health and the child's height-per-age Z score (HAZ) is the health outcome measure. Let child health, ***H ,*** be measured by child HAZ score. Because HAZ score is a cumulative measure of current and past nutritional status, HAZ score represents both $\boldsymbol{H_i}$ and $\boldsymbol{H_{io}}$, so that the equation can be modified as:

$$HAZ_i = \mathrm{H}(I_f, E_i, E_f, E_c, \theta) \qquad (3)$$

Based on a similar reduced form model,[5] Ponce, Gertler and Glewwe (1998) estimate that income has a small effect (.20-.28 change in HAZ score given a 1% change in income) on improving child growth. While their analysis argues that it would take a "very long time" for economic growth to improve Viet Nam's child nutrition situation, it does not convey if children, within families, are better or worse off after the reforms.

5.2 Identification of market openness

To measure the impact of market openness on child growth, the static household production model must be extended. Panel data, following families before and after the reforms, is the only type of data that allows us to determine the gainers and losers within households. Assume household panel data is available right before the reforms in 1987 and say five years after. We modify equation [3] by identifying the equation with a market openness variable, of $post87$, that takes on a value of 1 if the child was born after 1987 and 0 otherwise*:*

$$HAZ_{it} = \mathrm{H}(I_{ft}, post87_i, E_{ct}E_{ft}, E_{it}, \theta_t, \mu) \qquad (4)$$

where t indexes the year of the survey and the error term is partitioned into two components: θ_t and μ denoting time-varying and time invariant unobservables. Time-varying unobservables, θ_t, are characteristics that change over time such as better assimilation of health information by the mother with subsequent pregnancies. Time invariant unobservables, μ, would be characteristics of the household that are "fixed" over time such as religious or traditional practices in child rearing. Taking first-differences would purge the model of these "fixed effects" and can be expressed as:

$$\Delta HAZ_{it} = \mathrm{H}(\Delta I_{ft}, post87_i, \sum \Delta E_{ct} \sum \Delta E_{ft}, \sum \Delta E_{it}, \Delta\theta_t) \qquad (5)$$

From equation [5], we can construct a hypothesis test on the coefficient for $post87$.

5.3 Constructing the pseudo-panel

The objective is to measure if there are differences in the nutrition situation between two cohorts: children born in 1987 and before ($post87 = 0$) and children born after 1987 ($post87 = 1$). So far, we have assumed that panel data is available. Yet for most developing countries, including Viet Nam, this is not the case. [6] In the absence of panel data, there are several features of the VNLSS that can be empirically exploited. The sample of children under ten years old can be split in a comparably sized sample of children belonging in the post87=0 cohort and the post87=1 cohort. Moreover, if HAZ scores are used, human growth theory suggests that these two cohorts are comparable over different stages in the growth cycle for the following reasons:

1. Z-scores allow for comparisons between age groups for pre-pubescent children. It is not uncommon to compare the nutritional status of an infant with a ten-year-old using Z-scores (WHO 1995).
2. The most critical stage of linear growth for the child is from the prenatal period to two years of age (Martorell and Habicht 1986). Though growth continues during childhood, it is driven by the growth hormone and less determined by nutrition. Growth faltering after the age of two is an extension of the deficiencies in the infancy stage and the delayed initiation of the childhood phase of growth, supported by the growth hormone (Johnston 1986; Karlberg et al. 1994). The HAZ score (for pre-pubescent children) largely signals the growth faltering due to shocks during the child's time in the

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[5] The authors estimate an IV with FE specification identifying income with assets and non-wage income.
[6] A second round of the VNLSS has been fielded, but is not yet publicly available.

womb and first two years of life. In essence, we would be comparing if pre-reform children were better off when they were two years old than post-reform children at two years old.

For each family, the observations of each child be considered as a repeating observation, i such that:

$$HAZ_i = \mathrm{H}(I_i, post87_i, E_c E_f, E_i, \sigma_f, \nu) \tag{6}$$

The pseudo panel data allows for the partitioning of the error term into two components: denoting family-varying , σ_f and family invariant unobservables, ν . Family-varying unobservables would be individual child characteristics that change within families while family invariant unobservables, would represent characteristics that are "fixed" or idiosyncratic to the family. The model we have generated is a pseudo-panel approach for estimating policy induced changes. We can construct a hypothesis test for the *post87* variable and from this, infer if the impact of the reforms on child malnutrition has been positive or negative. As will be discussed in the next section, income is indexed for the individual child, rather than family income.

6. ESTIMATION ISSUES

In the pseudo panel approach, we are testing the period effects of cohorts. Do events after 1987 influence the nutritional outcome for families? Yet cohorts come with two other influences captured by the passage of time: lifecycle effects and age effects (Singleton, Straits and Straits 1993) that need to be isolated from the effect of market openness.

The lifecycle effect in this study context is principally the upward trend of family income through time. It could be that younger siblings are better off than their older brothers and sisters because they were born to older parents with more education, job experience and greater accumulated wealth. Hence if this upward trend in income is not controlled for, then we may be overestimating the effect of market openness on child growth. We control for some of these life cycle issues by identifying income with father's age (see Appendix 1). If income is identified as an estimated value of income at the time the child was born, then there would be unique income observations per child within a family. The equation to be estimated would be:

$$HAZ_i = \mathrm{H}(\hat{I}_i, post87_i, E_c, E_f, E_i, \sigma_f, \nu) \tag{7}$$

The lifecycle concern also falls under the empirical issue of unobserved heterogeneity within families. Over the lifecycle, the way that families respond to macroeconomic shocks, structural adjustment and regime changes may be idiosyncratic and fixed over time. We control for the income effect and other unmeasured idiosyncratic family characteristics (like the degree of entrepreneurial spirit of parents which makes them more successful in an open market regime, household) by taking family fixed effects (equation 8).

$$\Delta HAZ_i = \mathrm{H}(\Delta\hat{I}_i, post87_i, \Delta E_i, \nu) \tag{8}$$

The age effect needs to be disentangled because of the secular trend of HAZ scores as Vietnamese children grow older. We compare the mean height-for-age indicator of the Vietnamese sample with the mean height-for-age indicator of the NCHS/WHO reference population. The Vietnamese sample displays patterns of stunting commonly found in developing countries. Children start out with little difference from the healthy reference group, mean HAZ scores worsen as they get older, particularly for ages 19-24 months, after which it plateaus at a high level (Figure 1). If the $post87$ cohort consists of a larger proportion of older children, then the coefficient for $post87$ would be downward biased. On the other hand, the older children observed may be the "fittest" as they have survived the critical growth years. Thus, they may be selectively and thus systematically better off than younger children. Variations associated with being older are difficult to purge, but can be partially addressed with inclusion of categorical age variables. We include age categories corresponding to biological milestones in child linear growth. We spline age into 9 categories: 0 to 6 months, 7 to 12 months, 13 to 18 months, 19-24 months, 25 to 36 months, 37 to 48 months, 49 to 72 months, 73 to 96 months and 97 to 119 months. In addition, for a more flexible specification we interact gender with these age categories.

Other control variables include a dummy variable for oldest child since parental provision of resources towards a child may differ depending on the birth order of the child (Horton 1988). Age of mother at childbirth has also been shown to influence child health where risks of low birthweight or congenital abnormalities are highest for the youngest and the oldest of mothers. Children born to relatively young mothers are more likely to be underweight and to suffer from other health problems than are children born to older mothers (Gibson 1990).

We chose to construct dummy variables for not just the *post87* cohort of children, but also for the cohorts of children born after 1986, 1987, 1988, and 1990. We did this to account for the ups and downs of the reform process and to detect the marginal effects of each year after the *doi moi*. In 1988, the decollectivization of agriculture may have benefited children's nutritional status, but the 1989 privatization of the health sector may have been detrimental. Thus, we included dummy variables for each year after the reforms: *post86, post87, post88, post89, and post90*.[7] The comparison group would be the children born before *doi moi* in 1986. We can test the marginal effect of each year by examining the beta coefficient of each of the *post86, post87, post88, post89, post90* variables.

While the beta-coefficients measure the marginal effect of each year after the reforms, the F-statistic for these dummy variables would measure the cumulative effect. If the F-statistic for the cumulative dummy variables ($post86, post87, post88, post89, post90$) is negative, then the inference is that children born after 1986 are worse off in nutritional status than their siblings who were born before the reforms. The opposite would be inferred if the coefficient is positive. Finally, since there may be structural differences between the urban and rural sector, we ran the models for three stratifications: pooled, urban and rural.[8]

7. RESULTS

It is important to emphasize that the pseudo panel approach attempts to measure the nutrition gap of children within families. Taking a fixed effects specification within families purges the effect of family income from the model yet also swept away important household and community-level variables presented in Table 1 (see Appendix 2 for full model results). What is left on the right hand-side are the estimated income variable, mother's age at childbirth, oldest child, and the age and age-sex splines.

Table 3 presents the beta coefficients, t-statistics and F-statistics for the pooled sample. Model A tests the cumulative effect of the *post87* variable and Model B tests marginal effect of each of the post-reform variables. Looking at Model A, the cumulative effect of post87 is negative and significant (beta= -0.1803; t-statistic=-2.831). At the mean HAZ score of –2.01, those born after 1987 have HAZ scores that are lower by 0.1803 units than those born before 1987. This means that the market openness may have penalized the child by lowering HAZ scores (by 9%), but not to the extent of putting a child from a moderate to a severely malnourished state.

In Model B, we see that all of the post- *doi moi* variables are negative, but only the *post87* variable is significant at the 5% level. The *post86* and *post89* variables are significant at the 10% level. Model B's results generally correspond to the timing of policies. *Doi Moi* was announced in December 1986, so

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[7] We did not include dummy variables for post91 and post92 because the sample size would be too small and would include only children under 2.

[8] We constructed an F-statistic (Chow-test) to test the null hypothesis of no structural difference between the restricted (urban and rural) models with the unrestricted (pooled) model. We cannot reject that there are no structural differences between urban and rural areas: F = 6.85. Thus, separate regressions were run for the urban and rural samples.

differences for *post86* are not expected. Liberalization policies began in 1987 along with the government dropping food subsidies. But in 1988, the beginning of the decollectivization of agricultural land gave immediate benefits to net food producers, a majority of the rural sector. In 1989, some of these gains may have been offset by changes in the financing and delivery of care in the health sector and other effects of market openness. But by 1990, the differences between the *post90* children and the *pre1986* comparison group is virtually zero and insignificant.

Since eighty-five percent of the pooled sample are rural children, its results may obscure the direction of the impact of market transition on the urban sector. We turn now to the findings for the urban sample, and conclude with findings for the rural sample.

None of the post-reform variables were significant for the urban sample (Table 4), although for both Model A and B, the *post87* variable is positive. The sample size(n= 859) may be too small to estimate deviations from the mean. It is difficult to infer if market transition has penalized or rewarded the urban sector. For the rural sample, (Table5) the coefficient of the *post87* variable is larger (-0.228) and more significant (t-statistic = -3.271) than the coefficient of the *post87* for the pooled sample. And similar to the pooled sample, only the post87 variable is significant at the 5% level. Model B conveys that the rural sector faced a big penalty after 1987, but "recovered" after 1988 when food markets opened up. HAZ scores dipped after the privatization of health markets in 1989, but after 1990, the results indicate no effect of market openness on child growth.

8. DISCUSSION

The study's results provide evidence of a J-curve response of households to increasing market openness. At the initial stage, we find that market openness penalized child growth by a 9% reduction in HAZ scores. Through the course of transition, households are generally moving upward from the dip off the J-curve, though with some fluctuations. Households responded negatively to economy-wide "shock therapy" (Lipton and Sachs 1991) in 1987 but recovered somewhat with the opening of food markets in 1988. Child growth dipped again after the 1989 health sector reforms. But by 1991, market openness seemed to have a negative, near zero effect on child growth.

The study has several empirical and policy implications. First, the pseudo-panel approach we developed on this data yielded estimates that plausibly captured the timing of different market openness policies. Second, we did not expect that height-per-age, a stock indicator of nutritional status, would respond measurably to macroeconomic policy shocks in a short period of time. Indeed, past empirical work proves that the calorie-income elasticity is high for poor countries like Viet Nam (see Section 3) , but there is less evidence that health stock may actually erode (or improve) quickly in response to macroeconomic shocks. My third point is a corollary of the second point—children are a vulnerable population whose linear growth potential are affected by market openness reforms.

Notwithstanding the gains for consumers in the opening of markets for foods and other commodities, Viet Nam's transition to the market has led to more individualistic approaches in producing a healthy child. While households have the right to govern these decisions, the informational problems that arise, particularly during this period of transition, may be a source of increasing health status inequality. Governments embracing market openness should weigh the social cost of their policies on this group and consider taking back some of its responsibility from the market.

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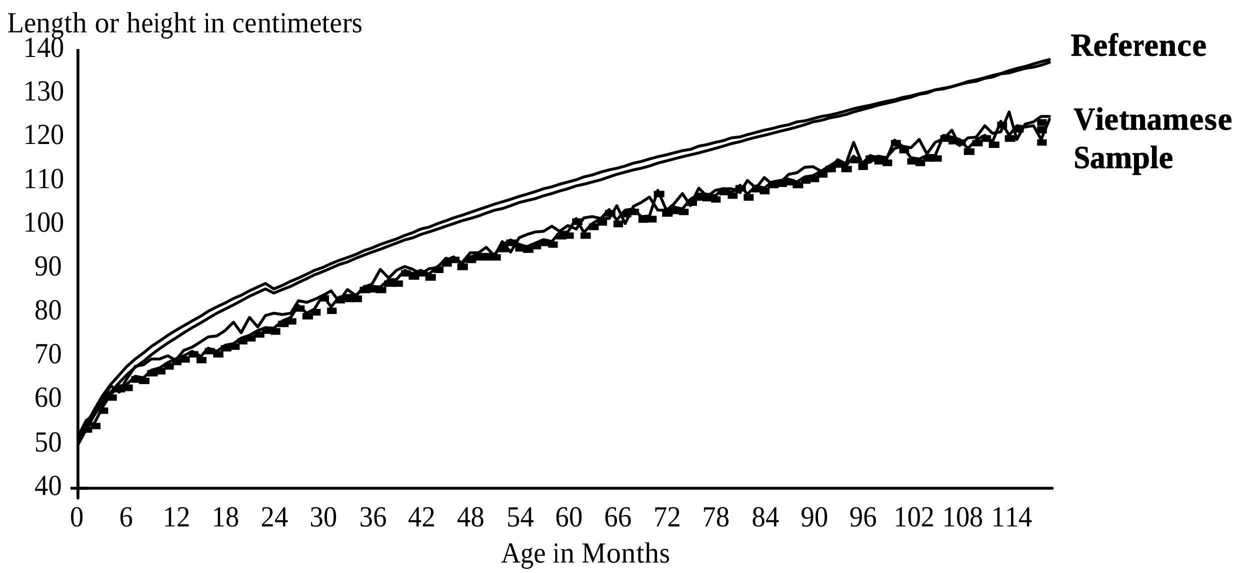
Figure 1: Comparison between NCHS and Vietnamese Median Heights[9]



[9]Note that the kink at 24 months for the NCHS sample is due to the change in the measurement of height. Children under 24 months are measured lying down (length) and children 24 months and older are measured standing up (height).

Table 1: Means for Community, Family and Child Characteristics

Community			Family			Child		
Indicators	mean	SD	Indicators	mean	SD	Indicators	mean	SD
Region			Mother's Height (cm)	151.8	4.96	Age in Months		
Northern Uplands	.18	n/a	Father's Height (cm)	161.9	5.50	0 - 6	.06	n/a
Red River Delta	.22	n/a	Mother's Education (years)	6.6	3.43	7 - 12	.05	n/a
North Central	.14	n/a	Father's Education (years)	7.5	3.17	13 - 18	.04	n/a
Central Coast	.11	n/a	Maternal Age (years)	27.4	6.03	19 - 24	.05	n/a
Central Highlands	.04	n/a	Ethnic Group			25 - 36	.10	n/a
Southeast	.12	n/a	Kinh	.83	n/a	37 - 48	.10	n/a
Mekong River Delta	.20	n/a	Tay	.02	n/a	49 - 72	.20	n/a
Rural	.85	n/a	Thai	.01	n/a	73 - 96	.20	n/a
Urban	.15	n/a	Chinese	.02	n/a	97 - 119	.20	n/a
			Khome	.02	n/a	Female	.51	n/a
			Muong	.05	n/a	Oldest Child	.30	n/a
			Religion of Head					
			Buddhist	.26	n/a			
			Christian	.09	n/a			
			Animist/Traditional	.01	n/a			
			Other	.02	n/a			
			None	.63	n/a			

Table 2: Mean HAZ-scores for Community, Family and Child Characteristics

Community		Family		Child	
Indicators	mean HAZ	Indicators	mean HAZ	Indicators	mean HAZ
Region		Mother's Height (at mean = 151.8)	-2.37	Age in Months	
Northern Uplands	-2.28	Father's Height (at mean = 161.9	-2.38	0 - 6	-0.51
Red River Delta	-2.04	Mother's Education		7 - 12	-1.40
North Central	-2.34	no schooling	-2.47	13 - 18	-2.04
Central Coast	-1.92	some primary	-2.19	19 - 24	-2.39
Central Highlands	-2.22	finished primary	-2.41	25 - 36	-2.10
Southeast	-1.55	finished lower secondary	-2.43	37 - 48	-2.15
Mekong River Delta	-1.84	finished upper secondary	-2.12	49 - 72	-2.24
Rural	-2.12	Father's Education		73 - 96	-2.17
Urban	-1.58	no schooling	-2.74	97 - 119	-2.07
		some primary	-2.36	Male	-2.09
		finished primary	-2.22	Female	-1.98
		finished lower secondary	-2.39	Female*Age Category	
		finished upper secondary	-2.22	0 - 6	-1.14
		Maternal Age (at mean = 27.4)	-2.37	7 - 12	-1.87
		Ethnic Group		13 - 18	-2.54
		Kinh	-1.98	19 - 24	-2.83
		Tay	-2.58	25 - 36	-2.33
		Thai	-1.91	37 - 48	-2.26
		Chinese	-1.59	49 - 72	-2.63
		Khome	-2.40	73 - 96	-2.36
		Muong	-2.15	97 - 119	-2.26
		Nung	-2.39	Not Oldest Child	-2.08
		Other	-2.42	Oldest Child	-1.95
		Religion of Head			
		Buddhist	-1.96		
		Christian	-2.14		
		Animist/Traditional	-2.67		
		Other	-1.69		
		None	-2.06		

Table 3: Pseudo-Panel Results: Pooled Sample

n = 5560	Household Fixed Effects			
	Model A		Model B	
Adj. R-squared	0.404		0.404	
Variables	beta	t-stats	beta	t-stats
*ln(Income)**	1.754♦	2.850	1.845♦	2.99
post86			-0.146	-1.86
post87	-0.1803♦	-2.831	-0.161♦	-2.45
post88			-0.118	-1.27
post89			-0.168	-1.80
post90			-0.065	-0.59
Mother's Age at Childbirth	.0247♦	2.227	0.026♦	2.31
Oldest Child	0.153♦	3.901	0.151♦	3.84
Constant	-15.243♦	-3.585	-15.450♦	-3.63
	F-Statistics of Joint Significance			
post86-post90 (5)			3.16♦	
Age Category (8)	12.66♦		12.16♦	
Females by Age Category (9)	2.96♦		2.98♦	
Degrees of Freedom	2597		2593	

♦ **Probability> t $\leq$ .05; Probability> F $\leq$ .05**

Income was identified by $dadag, dadage^2, dadadge^3, u$, where *dadage* is the variable name for father's age at the time the child was born, and *u*, the residual. The income variable is measured as real per capita expenditures, deflated to 1982 prices.

Table 4: Pseudo-Panel Results: Urban Sample

n = 859	Household Fixed Effects			
	Model A		Model B	
Adj. R-squared	0.383		0.397	
Variables	beta	t-stats	beta	t-stats
*ln(Income)**	-2.13	-1.006	-1.920	-0.89
post86			-0.256	-1.41
post87	0.1035	0.714	0.142	0.95
post88			-0.311	-1.28
post89			-0.410	-1.04
post90			0.039	0.16
Mother's Age at Childbirth	-.0077	-.409	-0.006	-0.31
Oldest Child	0.235♦	2.597	0.229♦	2.51
Constant	14.261	0.897	13.543	0.84

	F-Statistics of Joint Significance	
post86-post90 (5)		.98
Age Category (8)	5.01♦	5.00♦
Females by Age Category (9)	1.42	1.47
Degrees of Freedom	309	305

♦ Probability> t ≤ .05; Probability> F ≤ .05

Income was identified by $dadag, dadage^2, dadadge^3, u$, where *dadage* is the variable name for father's age at the time the child was born, and *u*, the residual. The income variable is measured as real per capita expenditures, deflated to 1982 prices.

Table 5: Pseudo-Panel Results: Rural Sample

n = 4701	Household Fixed Effects			
	Model A		Model B	
Adj. R-squared	0.396		0.395	
Variables	beta	t-stats	beta	t-stats
*ln(Income)**	1.85♦	2.835	1.917♦	2.93
post86			-0.127	-1.49
post87	-0.228♦	-3.271	-0.216♦	-3.00
post88			-0.073	-0.72
post89			-0.118	-1.23
post90			-0.065	-0.54
Mother's Age at Childbirth	0.031♦	2.574	0.032	2.63
Oldest Child	0.147♦	3.421	0.145	3.38
Constant	-15.98♦	-3.597	-16.11	-3.62

	F-Statistics of Joint Significance	
post86-post90 (5)		2.93
Age Category (8)	9.91	9.91
Females by Age Category (9)	3.00	2.96
Degrees of Freedom	2267	2263

♦ **Probability> t ≤ .05; Probability> F ≤ .05**

*Income was identified by $dadag, dadage^2, dadadge^3, u$, where *dadage* is the variable name for father's age at the time the child was born, and *u*, the residual. The income variable is measured as real per capita expenditures, deflated to 1982 prices.

Appendix 1 : Life-Cycle Effects of Income

The major life-cycle effect that may affect the household's food security and hence the child's nutritional status is parental income. If post-87 children are better off nutritionally, the policy implications are less clear. It could be that younger siblings are better off because they were born when their parents were older, had more education, longer job experience and more wealth. We control for some of these life cycle issues by controlling for father's age (a proxy of the age pattern of income).

The first step is to identify the father's age pattern of income. The intuition behind this is that income increases with father's age, peaks, then declines. We explore this relationship by identifying income with father's age. Real per capita household income $\left(\frac{\text{expenditures}}{\text{hhsize}}\right)$ is a function of father's age. We derived the functional form by saturating the right hand side of the equation with monotonic transformations of father's age (*dadage*, *dadage*^{2} and *dadage*^{3}).

$$\left(\frac{\text{expenditures}}{\text{hhsize}}\right) \quad = \quad f\left[dadage, dadage^2, dadage^3, u\right] \qquad [1]$$

From equation [1], let $(\overset{\wedge}{realeq})$ be the predicted value of income and $\hat{u}$ be the predicted value of the residuals. Since we know $(\overset{\wedge}{realeq})$ and $\hat{u}$, we can determine

$$f\left[dadage, dadage^2, dadage^3, u\right]. \qquad [2]$$

$$f\left[dadage, dadage^2, dadage^3, u\right] = \hat{realeq} + \hat{u} \qquad [3]$$

The left-hand side of equation [27] represents the age pattern of income. Let us call this *eqhat* so that

$$eqhat \ = \ (\overset{\wedge}{realeq}) \ + \ \hat{u} \qquad [4]$$

We include the natural log of *eqhat* as part of the right-hand side variables. We are able to assign an income variable for each child, so that taking family-level fixed effects will not sweep away the income coefficient.

Appendix 2: Pseudo-Time Series Study Model Results

	POOLED		URBAN		RURAL	
N		5560		859		401
Adj. R-squared		0.404		0.381		0.395
HAZ	Beta	t-stat	Beta	t-stat	Beta	t-stat
ln(per capita hh expenditures) identified by age of father at child's birth	1.845	2.99	-1.920	-0.89	1.917	2.93
post86	-0.146	-1.86	-0.256	-1.41	-0.127	-1.49
post87	-0.161	-2.45	0.142	0.95	-0.216	-3.00
post88	-0.118	-1.27	-0.311	-1.28	-0.073	-0.72
post89	-0.168	-1.80	-0.410	-1.04	-0.118	-1.23
post90	-0.065	-0.59	0.039	0.16	-0.065	-0.54
Age Category						
0-6	1.615	7.83	2.043	5.13	1.557	6.92
7-12	0.686	3.75	0.327	0.85	0.755	3.73
13-18	0.203	1.06	-0.209	-0.54	0.268	1.28
25-36	0.262	1.65	0.387	1.30	0.253	1.42
37-48	-0.036	-0.21	-0.018	-0.04	-0.019	-0.10
49-72	-0.219	-1.07	-0.631	-1.13	-0.139	-0.63
73-96	-0.461	-2.05	-0.739	-1.28	-0.400	-1.63
97-119	-0.315	-1.37	-0.838	-1.44	-0.228	-0.91
Female by Age Category						
0-6	0.199	0.96	0.215	0.52	0.182	0.82
7-12	0.310	1.44	1.665	3.28	0.108	0.46
13-18	-0.115	-0.62	-0.410	-0.93	-0.045	-0.23
19-24	-0.224	-1.32	-0.176	-0.45	-0.235	-1.25
25-36	-0.058	-0.58	0.048	0.22	-0.088	-0.79
37-48	0.011	0.12	-0.075	-0.20	0.005	0.05
49-72	0.013	0.21	0.117	0.81	0.009	0.14
73-96	0.177	2.93	-0.052	-0.36	0.202	3.07
97-119	0.200	3.26	-0.077	-0.43	0.239	3.64
mother's age at childbirth	0.026	2.31	-0.006	-0.31	0.032	2.63
oldest	0.151	3.84	0.229	2.51	0.145	3.38
constant	-15.450	-3.63	13.543	0.84	-16.111	-3.62

Speaker’s Biography

Ninez Ponce is an Assistant Professor in the Department of Health Services, UCLA School of Public Health and Senior Researcher at the UCLA Center for Health Policy Research. She is also a research consultant at RAND.

Dr. Ponce's current research interests include developing innovative approaches to expand insurance markets in developing countries and among low-income populations in the United States. Her research focuses on child health issues and access to care among linguistic and ethnic minorities in the United States. She is a co-Principal Investigator for an upcoming 55,000 household population health survey in California. Prior to joining the Center and the UCLA Public Health faculty, she was Deputy Director and Survey Research Manager for the Asian & Pacific Islander American Health Forum. Most recently, she served as RAND's resident Policy Adviser on health insurance reforms to the Ministry of Health, Republic of Macedonia. Trained in health economics and empirical modeling of policy issues, Dr. Ponce has a B.S.from the University of California, Berkeley, an MPP from Harvard University and a PhD from UCLA.