Crises, Food Prices, and the Income Elasticity of Micronutrients: Estimates from Indonesia

Emmanuel Skoufias, Sailesh Tiwari, and Hassan Zaman

The 2008 global food price crisis and more recent food price spikes have led to a greater focus on policies and programs to cushion the effects of such shocks on poverty and malnutrition. Analysis of the income elasticities of micronutrients and their changes during food price crises can shed light on the potential effectiveness of cash transfer and nutrition supplement programs. This article examines these issues using data from two cross-sectional household surveys in Indonesia, taken before (1996) and soon after (1999) the 1997–98 economic crisis, which led to a sharp increase in food prices. First, using nonparametric and regression methods, the article examines how the income elasticity of calories from starchy staples as a share of total calories differs between the two survey rounds. Second, the article estimates income elasticities of important nutrients in Indonesia. The analysis finds that, although summary measures such as the income elasticity of the starchy staple ratio might not change during crises, this stability masks important differences across individual nutrients. In particular, income elasticities of some key micronutrients, such as iron, calcium, and vitamin B1, are significantly higher in a crisis year than in a normal year, yet the income elasticities of others—such as vitamin C—remain close to zero. These results suggest that cash transfer programs might be even more effective during crises to ensure the consumption of essential micronutrients. But to ensure that all key micronutrients are consumed, nutrition supplement programs are also likely required.

JEL classification codes: I12, O12, D12, E31 keywords: Income elasticity, crisis, prices, micronutrients, vitamin C, starchy staple ratio

Global food prices rose sharply in 2008 and, in real terms, reached levels not seen since the early 1970s. The Food and Agriculture Organization (FAO 2008) Food Price Index grew by 73 percent between September 2006 and mid-2008, driven by unprecedented increases across all food categories. During the same period meat and fish prices increased 25 percent, eggs and milk 91 percent, and vegetables, fruits, and sugar 61 percent. This study examines the changes in the income elasticity of micronutrients during a food price crisis. The analysis finds that, although summary measures such as the income elasticity of the starchy staple ratio might not change during crises, this stability masks important differences across individual nutrients. In particular, income elasticities of some key micronutrients, such as iron, calcium, and vitamin B1, are significantly higher in a crisis year than in a normal year, yet the income elasticities of others—such as vitamin C—remain close to zero. These results suggest that cash transfer programs might be even more effective during crises to ensure the consumption of essential micronutrients. But to ensure that all key micronutrients are consumed, nutrition supplement programs are also likely required.

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percent, oils and fats 149 percent, and cereal grains 123 percent. After a downward trend in 2009, global food prices rose again in late 2010, with prices in December 2010 close to their 2008 peak. Such food price volatility, particularly sharp spikes, needs to be monitored closely for its impact on the poor. A review of the literature on the effects of the 2008 food price increases suggests that they are likely to have had a significant impact on the incidence of poverty (Ivanic and Martin 2008) and undernourishment (Tiwari and Zaman 2010) throughout the developing world.

Soaring food prices and their adverse effects have not only heightened concerns about food security and malnutrition in parts of the developing world, but have also sparked a renewed interest in the design of effective policy responses. From the household point of view, such price increases have two main consequences: they reduce the purchasing power of household income, especially among poorer households, which spend a larger share of their incomes on food, and they result in a relative price effect that induces households to substitute away from more expensive foods.¹

Government interventions during rising food prices have almost always been motivated by the need to compensate poor households for their lost purchasing power. These interventions—commonly known as social safety net programs—are aimed at smoothing consumption and protecting the caloric availability of households to prevent increases in poverty and hunger. A review of the safety net programs used during the 2008 food price crisis shows that they typically took the form of income support, cash transfers, price subsidies, and supplementary feeding programs or in-kind transfers of staple foods (Wodon and Zaman 2010).

Yet even if income support or in-kind staple food distribution is successful at preventing available calories from reaching dangerously low levels, there are valid concerns about dietary diversity and the consequent risk of malnutrition. When household income drops, households may keep calorie levels more or less constant through substitutions within and between food groups, while the consumption of essential micronutrients might decrease significantly as households consume less meat, vegetables, eggs, and milk (Behrman 1995). The extent to which the consumption of micronutrients responds to decreases in income among poor households is of particular concern given the long-run consequences that a diet poor in micronutrients can have on child development before and after birth.

Though there is ample evidence on the income elasticity of calories (Strauss and Thomas 1995; Subramanian and Deaton 1996; Skoufias 2003), empirical evidence on the income elasticity of micronutrients is sparse. The evidence that does exist suggests substantial differences in the income elasticities of

¹. Purchasing power improves for net sellers of agricultural commodities whose prices increase. In addition, for households close to subsistence and already consuming the cheapest sources of calories, substitution possibilities are much more limited.
micronutrients (Behrman and Deolalikar 1987; Bouis 1991). Pitt and Rosenzweig (1985), focusing on Indonesia and using data from farm households, report very low income elasticities (less than 0.03) for a set of nutrients that included calories, proteins, fats, carbohydrates, calcium, phosphorus, iron, and vitamins A (carotene) and C (ascorbic acid). Chernichovsky and Meesook (1984), using data from rural and urban areas, report much higher income elasticities for nutrients—for example, from 0.7 to 1.2 for the poorest 40 percent (by expenditure) of the population on Java. Similarly diverse estimates have been reported for other countries.2

Most of the empirical evidence sheds light on whether the price sensitivity of the demand for food and nutrients varies with income level (Timmer and Alderman 1979; Timmer 1981; Pitt 1983) or whether the income elasticity of calories varies with income level (Behrman and Deolalikar 1987; Ravallion 1990; Strauss and Thomas 1995; Subramanian and Deaton 1996). But evidence is lacking on whether the income elasticity of nutrients varies significantly depending on changes in the relative prices faced by households, such as changes experienced during food price spikes. This is an important gap in the literature and perhaps a large contributor to the significant divergence in estimates of the income elasticity of nutrients. When income elasticities are estimated using cross-sectional data, price variations necessarily come from the cross-section as well. But despite being pointed out by Deaton and Muellbauer (1980), the possibility that these estimated elasticities could be sensitive to the degree of variability of prevailing relative prices is something that has neither been tested explicitly nor been used to qualify the policy recommendations that emerge from existing studies.

In 1998, during the financial crisis in Indonesia, the value of the rupiah depreciated dramatically, falling from around 2,400 per US dollar in June 1997 to just under 15,000 per dollar in June 1998 and finally settling at 8,000-9,000 per dollar in December 1998. These fluctuations in the exchange rate led to large increases in the prices of tradable commodities in domestic markets. Indonesia’s consumer price index rose 107 percent between February 1996 and February 1999. During the same period the food price index jumped 188 percent. In addition, subsidies for consumer goods such as rice, oil, and fuel were removed in 1998. So it is questionable whether estimates of the income elasticity of nutrients obtained from a sample of households observed during

2. Behrman and Deolalikar (1987), using data from villages covered by the International Crops Research Institute for the Semi-Arid Tropics (ICRISAT), report income elasticity estimates of 0.06 to 0.19 for proteins (depending on whether level estimates or differences over time are used), 0.30 to –0.22 for calcium, –0.11 to 0.30 for iron, 0.19 to 2.01 for vitamin A, –0.08 to 0.18 for thiamine, 0.69 to 0.01 for riboflavin, –0.15 to 0.21 for niacin, and 0.15 to 1.25 for vitamin C. A Nicaraguan study by Behrman and Wolfe (1984) reports significant income elasticity estimates in the range of 0.04 to 0.11 for calories, proteins, iron, and vitamin A (with statistically significant, but quantitatively small, nonlinearities). A Philippine study by Bouis (1991) reports an income elasticity of 0.44 for iron, an income elasticity of 0.16 for calories, and insignificant income elasticities for vitamin A and vitamin C.
precrisis years can provide any guidance on how caloric and micronutrient availability might respond to additional income (other things being equal) during a period with a different set of relative prices. From a policy perspective, the sensitivity of the income elasticity of nutrients to relative prices implies that policies aimed at increasing household income—such as employment and cash transfer programs—might be less effective in protecting nutritional outcomes under some economic conditions.

This article uses shocks to food prices in Indonesia to examine the relationship between nutrient consumption and prices, with the analysis conducted at two levels. First, using the starchy staple ratio (the calories from starchy staple foods such as cereals and tubers as a share of total calories) as a summary measure of household nutritional welfare, it assesses the impact of dramatic changes in food prices on household dietary composition. The SSR is defined as the share of calories consumed obtained from starchy staple foods such as cereals and tubers. According to Bennett’s Law this ratio is inversely related to the importance of starches relative to higher-quality, more expensive, micronutrient-rich foods such as meat, fish, fruits, and vegetables. The focus is on how the income elasticity of the starchy staple ratio differs between survey rounds in 1996 and 1999, when relative prices were very different for cereals and the other major food groups. The analysis is conducted for Indonesia’s entire population and for the poor in urban and rural Central Java (one of the country’s poorest provinces) in 1996 and 1999. Results are reported using both nonparametric and regression methods.

This analysis is supplemented by updated estimates of the income elasticities of important nutrients in Indonesia, such as calories, proteins, fats, carbohydrates, calcium, phosphorous, iron, and vitamins A, B1 (thiamin), and C. In times of crisis, cash transfers may be the fastest and least costly way of reaching households most likely to be adversely affected, if the delivery infrastructure is in place and leakage is low. Reliable elasticity estimates can help policymakers determine beforehand whether cash transfers are likely to increase nutrient availability among poor households or if other interventions may be needed. If estimated elasticities are high and significantly different from zero, policies such as cash transfers that aim to compensate for price increases are likely to be effective. But if estimated elasticities are indistinguishable from zero, alternative interventions may be needed. Particular emphasis is placed on the sensitivity of the elasticity estimates to biases due to measurement errors in consumption and nutrient availability at the household level.

The article also presents a test of whether the income elasticity of nutrients varies with the economic conditions facing households. Other things being

3. As a measure of dietary composition, the starchy staple ratio is useful because it summarizes changes in nutritional welfare over time. There is also evidence from Indonesia that a closely related variable—the share of total food expenditure going to nongrains, such as animal and plant sources—is positively associated with lower incidence of stunting among children younger than five (Sari and others 2010).
equal, changes in the relative prices of staple foods may result in unexpected responses of how the demand for nutrients responds to cash transfers. For example, if total calorie availability is already low and the price of a staple increases during a crisis, households receiving cash transfers might spend the additional income on that same staple if it continues to be the cheapest source of calories (Behrman 1988; Behrman and Deolalikar 1989).

This article is structured as follows. Section I describes the data used for the analysis and the construction of key variables and presents background information on the changes in prices and nutrient availability between 1996 and 1999 in Indonesia. Section II discusses the empirical strategy and the results of estimation using both nonparametric and regression methods. Section III summarizes the findings and presents some policy implications.

I. Background and Data

The analysis in this article is based on the detailed consumption module of the National Socio-Economic Survey (SUSENAS) conducted every three years by the Central Statistical Agency of the Government of Indonesia. The consumption module is nationally representative of urban and rural areas in each of the country’s 26 provinces and the Jakarta metropolitan area. The survey included 60,678 households in 1996 and 62,217 households in 1999.

In addition to the detailed nature of the survey, one of the main advantages of comparing the income elasticity of calories in these two years is the opportunity to examine economic behavior under dramatically different relative price regimes. In February 1999, when the 1999 SUSENAS was conducted, inflation in Indonesia had reached its peak since the start of the financial crisis in late 1997 and its intensification in mid-1998. An additional benefit is that the same questionnaire was used at the same point in time in both survey years. Thus the possible influence of seasonal factors in the relationship of income to calories—as emphasized by Behrman, Foster, and Rosenzweig (1997)—can be controlled for. A detailed discussion of the SUSENAS consumption module and the construction of key variables used in the analysis is presented in appendix A.

To make meaningful comparisons across two cross-sectional surveys that are three years apart, the nominal income of households in 1999 must be expressed

4. This statement is not intended to compare the effectiveness of cash transfers relative to other possible alternatives of increasing nutrient availability within households. Alternatives include in-kind food transfers and employment creation programs.

5. The analysis also uses some variables from the larger SUSENAS (core survey) containing observations for about 205,000 households.

6. The Idul Fitri-Lebaran holiday following the fasting month (Ramadan) is a moving holiday—in 1999 it fell in late January. Central Statistical Agency officials confirmed that the survey was conducted two weeks after the holiday, so the value of household food consumption has little chance of appearing unusually high due to the holiday.
in 1996 rupiah. A critical point for the construction of real income in 1999 is the fact that changes in food prices affect households differently depending on the share of their budgets spent on food. Typically, poorer households spend a much larger share of their incomes on food—nearly 60 percent for poor rural households in Indonesia, compared with 40 percent for urban households at the top of the expenditure scale.

The SUSENAS consumption module includes data on the value and quantity of food items consumed, allowing calculations of unit values at the household level. A deflator was constructed combining the unit values calculated from the consumption module with province-specific prices reported for nonfood items by the Central Statistical Agency. First, because expenditures are only collected for nonfood items, a deflator for nonfood items was constructed using the mean shares of major groups of nonfood items in the February 1999 survey as weights and the province-specific price indexes for these groups. Second, a household-specific food deflator was calculated from a weighted average of the 52 food items used in the calculation of the poverty line in Indonesia.

Specifically, the household-specific food deflator was calculated using the formula

\[
P_{F}(99) = \left[ \sum_{i=1}^{52} \left( s_{i}^{h}(99) P_{i}(R, 96) / P_{i}(R, 99) \right) \right]^{-1}
\]

which is the standard formula for calculating a Paasche price index (Deaton and Zaidi 1999). \( s_{i}^{h} \) denotes the share for food item \( i \) of the total amount expended on the 52 food items, and the superscript \( h \) indicates that the share varies by household. The second term is the ratio of the median unit value of food item \( i \) in region \( R \) in 1996 to the corresponding unit value in 1999. Household-specific unit values of food items are replaced by median unit values for each of the urban and rural areas of the 26 provinces and the Jakarta metropolitan area (a total of 53 regions) to minimize the influence of measurement errors and differences in the quality of food consumed by wealthier households (Deaton 1988). With these price deflators for food and nonfood, the overall price deflator for household \( h \) in 1999, \( P^{h}(99) \), can be expressed as

\[
P^{h}(99) = \tilde{W}_{F}^{h}(99) P_{F}^{h}(99) + \left( 1 - \tilde{W}_{F}^{h}(99) P_{NF}^{h}(R, 99) \right).
\]

Note that the weights applied to food and nonfood also vary across households. The weight for each household is calculated from the predicted value of

7. More details on the construction of the price indexes can be found in Skoufas (2003), Suryahadi and others (2000), and Levinsohn, Berry, and Friedman (1999), all of whom take a similar approach to constructing household-specific price indexes for Indonesia.

8. The province-specific price indexes for food and nonfood groups reported by the Central Statistical Agency are based on prices for 27 cities in 1996 and 44 cities in 1999.
the regression of household food share in 1999, \( \hat{W}_h(99) \), on the logarithm of per capita consumption and the logarithm of household size. This approach eliminates the influence of household-specific, unobserved components (such as taste preferences) on the share of food.

To provide more concrete evidence about the relative price regimes prevailing in the two survey years, figures 1a and 1b show the changes in mean prices per 1,000 calories (1 kilocalorie) paid by rural and urban households between 1996 and 1999 in Central Java, a densely populated province with a high concentration of poor people. Prices per kilocalorie are calculated by dividing the nominal value of household consumption for each food group by the

**Figure 1. Changes in Price of 1,000 Calories by Food Group Relative to Cereals and Tubers in Rural and Urban Central Java, 1996 and 1999**

Source: Authors’ calculations based on 1996 and 1999 SUSENAS consumption modules using expenditure quartiles.
kilocalories provided by all the items in the group. Poorer households may consume food items of lower quality and, as a result, the prices per kilocalorie paid by these households may be lower than those paid by richer households. To investigate for this possibility, prices per kilocalorie are calculated separately for households at the bottom and top 25 percent of the distribution of total consumption per capita in each year.

Figures 1a and 1b confirm that the relative prices faced by households changed considerably between 1996 and 1999. In general, the absolute prices of calories from all food groups seem to have increased dramatically between the two years in both urban and rural areas. Relative to cereals, food groups such as meat and fish, fruits and vegetables, and eggs and milk became more expensive in both rural and urban areas. There is also considerable heterogeneity between the rich and the poor in the magnitude of the relative price changes. Price increases appear more pronounced for the poor, particularly in urban areas. For example, in urban areas the price of eggs and milk relative to cereals increased 23 percent for the poorest consumers but only 4 percent for the richest consumers.

Price changes of this magnitude undoubtedly had a large impact on calorie and nutrient availability at the household level. Figures 2a and 2b present the change in average daily calories per capita between 1996 and 1999 for rural and urban Central Java. There was a significant reduction in both total and proportional calorie availability. Calorie availability at the household level in Central Java declined by 8 percent in rural areas and 6 percent in urban areas. There was a much larger reduction in calories derived from food groups richer in micronutrients—meat and fish, fruits and vegetables, and eggs and milk. The figure also shows considerable heterogeneity in the reduction between the rich and the poor, with larger reductions for the poor in almost all cases in both rural and urban Central Java.

The change in calories sourced from cereals and tubers is interesting. In rural Central Java calories obtained from cereals and tubers fell 12 percent among the poorest households, but only 1 percent among the richest households. By contrast, roughly the same level of decline was experienced by the poorest and richest households in urban areas. This discrepancy is largely because rich rural consumers are likely to be landowners, possibly engaged in the production of some of these staples. Thus price increases for these staples

9. The calorie prices reported are derived by dividing expenditures by total calories in the food group in each year. Thus the price of calories in 1999 may be biased downward depending on the extent to which households substituted away from more expensive food items within and between groups.

10. In 1999 the percentiles of total consumption per capita are estimated after dividing consumption by the deflator discussed earlier.

11. For a related analysis of the impact of the Indonesian crisis on budget shares, with repeated observations on sampled households, see Thomas and others (1999).
II. Empirical Framework, Analysis, and Results

Economic theory provides little guidance on how the income elasticity of a given commodity might change with changes in prices. Given a Marshallian demand function for any food item $x_i$, summarized by the function $x_i = x(p, M, Z)$ where $\hat{p}$ is a vector of relative prices, $M$ is real income, and $Z$ is a vector of preference shifters (such as household demographic characteristics), it follows that, in general, the response of demand to income changes depends

Figure 2. Changes in Per Capita Calorie Consumption by Food Group in Rural and Urban Central Java, 1996 and 1999

Source: Authors’ calculations based on 1996 and 1999 SUSENAS consumption modules using expenditure quartiles.

may have made those rich consumers better off and not affected their consumption of cereals and tubers by much.
on the same set of variables:

\[ \frac{\partial x_i}{\partial M} (\bar{p}, M, Z). \]

Without strong (even arbitrary) assumptions about the separability of preferences between and within specific food groups, little can be said (at least from a theoretical perspective) about how changes in prevailing relative prices might affect the demand for a commodity in response to income changes. Under these circumstances this issue can be addressed only empirically. This is precisely the gap that this article aims to fill. The question addressed empirically is whether income elasticity differs significantly between a noncrisis year (1996) and a crisis year (1999), which have very different relative price vectors, \( \bar{p}_{96} \) and \( \bar{p}_{99} \):

\[ \frac{\partial x_i}{\partial M} (\bar{p}_{96}, M, Z) \neq \frac{\partial x_i}{\partial M} (\bar{p}_{99}, M, Z) \] (3)

The analysis uses both nonparametric and regression methods that take into account the role of measurement error in total outlay.

The “almost ideal demand system” proposed by Deaton and Muellbauer (1980) provides the empirical framework for the analysis because it makes transparent the dependence of income elasticity on relative prices. Starting with the Working-Leser formulation that relates value shares to the logarithm of total expenditures

\[ w_i = \alpha_i + \beta_i \log(x) \] (4)

Deaton and Muellbauer propose a way of making the coefficient of total expenditures, \( \beta_i \), a function of prices. Given a cost function

\[ \log c(u, p) = a(p) + ub(p) \] (5)

and choosing the functions \( a(p) \) and \( b(p) \) to be of the form

\[ a(p) = \alpha_0 + \sum_k \alpha_k \log p_k + \frac{1}{2} \sum_k \sum_l \gamma_{kl} \log p_k \log p_l \] (6)

\[ b(p) = \beta_0 \Pi_k p_k^{\beta_k} \] (7)

the Engel curve can be expressed as

\[ w_i = \alpha_i + \sum_j \gamma_{ij} \log p_{ij} + \beta_i \log(x/p) \] (8)

12. The quadratic Engel curve proposed by Banks, Blundell, and Lewbel (1997) is an extension of the almost ideal demand system specification that also explicitly recognizes the dependence of the income elasticity to prices (see equation 11 in their paper).
where \( P \) is the price index defined by

\[
\log P = a_0 + \sum \alpha_k \log p_k + \frac{1}{2} \sum_k \sum_l \gamma_{kl} \log p_k \log p_l
\]  

and the parameters \( \gamma_{ij} \) are defined as

\[
\gamma_{ij} = \frac{1}{2} \left( \gamma_{ij}^* + \gamma_{ji}^* \right) = \gamma_{ji}.
\]

Based on this framework, the question addressed in this article translates to whether the \( b \)'s of equation (7), estimated using cross-sectional data from 1996—the normal year—as the coefficients of the logarithm of total expenditures in equation (8), are statistically and economically similar to those estimated in 1999—the crisis year.

**Nonparametric Regression and the Starchy Staple Ratio**

The analysis begins with an investigation of how the starchy staple ratio (SSR) and its income sensitivity vary in a cross-section of households between 1996 and 1999. SSR, calculated as the share of total calories obtained from cereals and tubers, is considered a more useful aggregate measure of household welfare than is total caloric availability per capita because SSR captures diversity in diets. According to Bennett’s Law, SSR declines with household income.\(^{13}\) This analysis adopts a flexible approach that examines Indonesian households by aggregating the caloric contents of the more than 200 food items included in the SUSENAS survey. Given that the income elasticity of calories in Indonesia is known to be nonlinear (Skoufias 2003), nonparametric methods are used that also allow the income elasticity of the SSR to vary with income level.

Using \( y \) to denote the logarithm of SSR and \( x \) the logarithm of per capita total household consumption, the regression function is:

\[
m(x) = E(y/x).
\]

Following Subramanian and Deaton (1996) and Deaton (1997), \( m(x) \) is estimated using a smooth local regression technique proposed by Fan (1993).\(^{14}\) At any given point \( x \), a weighted linear regression is run of the logarithm of SSR on the logarithm of per capita consumption. The weights are chosen to be the largest for sample points close to \( x \) and to decrease with distance from \( x \). The distribution of the logarithm of per capita consumption is divided into 100 evenly spaced grids, and local regressions are estimated for each grid instead of

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\(^{13}\) Bennett’s law involves the relation between household diet and income, while Engel’s law relates the share of food expenditures in a household budget to household income. Timmer, Falcon, and Pearson (1983) provide a detailed discussion of Bennett’s law.

\(^{14}\) Fan (1993) demonstrates the superiority of the smooth local regression technique over kernel and other methods.
for each point $x$ in the sample. For the local regression at $x$, observation $i$ has the quartic kernel weight

$$w_i(x) = \frac{15}{16} \left[ 1 - \left( \frac{x - x_i}{h} \right)^2 \right]^2$$

if $-h \leq x - x_i \leq h$ and zero otherwise. The quantity $h$ is a bandwidth set to trade off bias and variance, which tends to zero with increasing sample size. This analysis uses a bandwidth of 0.8.

Figures 3a and 3b compare logarithms of SSR of dietary content in 1996 and 1999 with logarithms of per capita expenditure (lnPCE), using 1996 prices, for rural and urban Central Java. With the income level in 1999 made comparable to that in 1996, these figures make it possible to examine the effects of changes in food prices between 1996 and 1999 on household dietary composition (as summarized by SSR), assuming that family size did not change significantly during that period.

An advantage of these nonparametric figures is that potential biases in the measurement of calorie availability among higher-income households do not affect estimates of nutrient income elasticity among poorer ones. For example, figures 3a and 3b make it possible to obtain a sense of the extent to which the exclusion of nutrients obtained from prepared foods (consumed more by wealthier households) affects the nutrient income elasticity of wealthier households.

In rural Central Java in 1999 SSR was just below the median level of per capita expenditure. SSR was slightly below its level in 1996 for households in the lower half of the distribution and above it for households in the upper half (see figure 3a). This suggests that in the crisis year the availability of calories from cheaper food sources (such as cereals and tubers) generally increased for higher-income groups in rural areas. This pattern is even more obvious in urban areas, where the 1999 SSR line lies almost uniformly above the 1996 SSR line. Thus for the urban poor there is a rather clear shift to the right in 1999, indicating a larger reliance on starchy staples for calories in times of crisis. Among rural households the pattern is less clear.

The differences in the slopes of the SSR lines in 1999 and 1996, in both rural and urban areas, suggest that the responsiveness of SSR to increases in income varies in crisis and noncrisis years (figures 3c and 3d). In rural areas the elasticity of SSR relative to income is higher during the crisis year and invariant to household income at about $-0.27$ percent. Thus it appears that an increase in income during a crisis year, such as cash transfers to poor rural households, is likely to be more effective in increasing dietary diversity (that is, lowering SSR) than in a noncrisis year.

15. Studdert, Frongillo, and Valois (2001) corroborate the increase in household food insecurity and compromised diet during the crisis in three Java provinces, including Central Java.
In contrast, in urban areas, where there was a larger reliance on starchy staples for calories in the crisis year, the income elasticity of SSR was about the same in 1996 and 1999 for poorer households, becoming smaller than the elasticity in the noncrisis year.\textsuperscript{16} Note, however, that as bivariate plots of calorie

\textbf{FIGURE 3. Features of the Starchy Staple Ratio in Urban and Rural Central Java, 1996 and 1999}

\textbf{Relationship between Starchy Staple Ratio and Per Capita Expenditure}

\begin{enumerate}
  \item \textbf{(a)}
  \begin{itemize}
    \item 1996
    \item 1999
  \end{itemize}
  \textbf{Rural}

  \begin{itemize}
    \item \textbf{InPCE}
    \item \textbf{InSSR}
  \end{itemize}

  \item \textbf{(b)}
  \begin{itemize}
    \item 1996
    \item 1999
  \end{itemize}
  \textbf{Urban}

\end{enumerate}

\textbf{Relationship between Income Elasticity of Starchy Staple Ratio and Per Capita Expenditure}

\begin{enumerate}
  \item \textbf{(c)}
  \begin{itemize}
    \item 1996
    \item 1999
  \end{itemize}
  \textbf{Rural}

  \begin{itemize}
    \item \textbf{InPCE}
    \item \textbf{Income elasticity of SSR}
  \end{itemize}

  \item \textbf{(d)}
  \begin{itemize}
    \item 1996
    \item 1999
  \end{itemize}
  \textbf{Urban}

  \begin{itemize}
    \item \textbf{InPCE}
    \item \textbf{Income elasticity of SSR}
  \end{itemize}

\textit{Source:} Authors’ calculations based on 1996 and 1999 SUSENAS consumption modules.

\\textsuperscript{16} Supplemental appendix S3, available at http://wber.oxfordjournals.org, examines whether the elasticity estimates are significantly different at different levels of outlay by checking whether the standard error bands for the 1996 estimate overlap those for the 1999 estimate. The figures do not reveal any significant differences in the income elasticity estimates for the crisis and noncrisis years, a result confirmed by the regression methods used in the latter part of this article.
shares at the household level these are descriptive only. The next section performs a more detailed analysis of SSR and its elasticity.

Regression Analysis

The analysis in the previous section focused on the bivariate relationship between SSR and per capita expenditure. Though being informative about the general shape of that relationship and how it changed between 1996 and 1999, that analysis cannot account for a number of critical factors—primarily the differences in the age and gender composition of households, as well as the problem of correlated errors between household consumption and nutrients. This section examines the elasticity of SSR relative to household consumption, controlling for a set of household characteristics.

The focus on the effects of relative price vectors on the income elasticity of demand faces some constraints. A typical cross-sectional household survey collects data within a short timeframe. As a result, most of the price variation for any household commodity comes from differences in transport costs, market segmentation, the quality of the commodity, and other transaction costs that can prevent the equalization of prices paid by consumers for the same commodity.

To the extent that households in different locations are surveyed at different times of year, a survey might also capture seasonal variability in prices. Even so, it is doubtful that seasonal variation in prices provides an adequate representation of the change in relative prices that consumers face during an economic crisis. Household panel data provide an opportunity to circumvent some of these shortcomings. Behrman and Deolalikar (1987), for example, analyze the relationship of calories to income using data from village surveys conducted by the International Crops Research Institute for the Semi-Arid Tropics (ICRISAT). But even these data shed little light on the relationship between household food consumption and spending during a crisis, because the economic environment was relatively stable during the period of that study.

To compensate for the fact that these are cross-sectional data from surveys taken in different years, not longitudinal data, this analysis uses a flexible specification that provides an explicit test of the difference between the elasticity coefficients in the precrisis and crisis years. To implement this, the two sets of cross-sectional data were pooled, then the following regression was run:

\[
\ln sSR_{jkt} = (\alpha_{96} + \beta_{96} \ln PCE_{jkt} + \gamma_{96} X_{jkt} + \mu_{96})
+ D_{99} \left( \alpha_{99} + \beta_{99} \ln PCE_{jkt} + \gamma_{99} X_{jkt} \right) + \mu_{99} + \varepsilon_{jkt}
\]

17. In much of the literature on the nutrient-income relationship (see Strauss and Thomas 1995), prices are typically left out of the specification of the Engel curve estimated using cross-sectional data. This is based on the ad hoc assumption that all households face the same prices. Notable exceptions are Subramanian and Deaton (1996), Behrman and Deolalikar (1987), Bouis and Haddad (1992), and Banks, Blundell, and Lewbel (1997).
where \( \ln SSR_{jkt} \) is the natural logarithm of SSR for household \( j \) that lives in cluster \( k \) in year \( t \), \( \mu_{96} \) and \( \mu_{99} \) are vectors of binary variables identifying the cluster fixed effects in the 1996 and 1999 rounds, and \( D_{99} \) is a binary dummy variable equal to one for observations in 1999 and zero otherwise. The variable denotes the natural logarithm of real per capita consumption expenditures for household \( j \) that lives in cluster \( k \) in year \( t \). \( e_{jkt} \) is an error term summarizing the influence of random disturbances. The set of control variables \( X_{jkt} \) is a vector of household characteristics indicating the logarithm of household size and variables characterizing the age and gender composition of the household, all expressed as ratios of the total family size (the number of children ages 0–5, the number of children ages 6–12, the number of male and female household members ages 13–19 and 20–54, and the number of men older than 55). Additional binary variables include whether the household head is a woman, dummy variables on the education levels of the household head and his or her spouse (completed primary, junior high, or senior high school), and the sectors of employment of the household head and his or her spouse (self-employed, unemployed, or a wage worker).

With this specification, cluster-level fixed effects are allowed to differ across the two survey years, providing control over differing relative prices between years. In addition, the coefficients of all the control variables and \( \ln PCE_{jkt} \) are allowed to vary across the two years, providing an explicit test of the difference in the income elasticity of the various dependent variables between 1996 and 1999. The dummy variable \( D_{99} \) is also able to absorb any other aggregate effects (aside from relative prices) that might have changed between 1996 and 1999.

As first pointed out by Bouis and Haddad (1992), a food expenditure survey can overstate the nutrient availability in wealthier households, since it is common for these households to provide meals to employees and domestic servants. In addition, following the 1997–98 crisis, it is plausible that there was an increase in the frequency of this practice. To minimize potential problems introduced by the fact that the level (and thus the elasticity) of SSR and nutrients may be less accurately measured for wealthier households, the estimation limits the sample to the lower half of the distribution of consumption per capita in rural and urban Central Java. Robust standard errors are estimated to control for unknown forms of heteroskedasticity.

18. The SUSENAS survey is a clustered survey with at most 16 households surveyed per cluster each year. Clusters in 1996 and 1999 have the same code, but it is unclear whether they represent the same clusters across the two years. Because of this limitation, clusters with the same code in different years are treated as different clusters.

19. Subramanian and Deaton (1996) use the same approach with cross-sectional data. Cluster-level fixed effects also take into account other time-invariant local characteristics that determine dietary intakes and preferences.

20. In the SUSENAS survey, domestic servants are counted as household members.
In defense of the ordinary least squares (OLS) estimates, it is important to bear in mind that the SUSENAS survey is a seven-day food intake and consumption survey that carefully collects information about food consumed outside the household as well as food received in kind from outside sources. The survey also captures food received as payment for services, food gifts, and food transfers. Since the same questionnaire was applied at the same point in time each survey year, there is no reason to believe that biases exist due to these factors.

Yet the possibility remains that correlated measurement errors in total food consumption (and thus calorie and nutrient availability) are potential sources of bias in estimates of income elasticities of nutrients. As first noted by Bouis and Haddad (1992) in a linear version of equation (13), the likelihood that measurement errors in nutrient availability are positively correlated with measurement errors in household consumption implies that this type of measurement error is not the standard errors-in-variables problem, where coefficients are likely to be biased toward zero. In the context of correlated measurement errors in the dependent and independent variables of a regression, it is unclear whether the upward bias from the correlated errors outweighs the standard downward attenuation bias from the measurement error in total consumption. So the direction of net bias in income elasticity estimates obtained using OLS methods will generally depend on the relative size of the correlation between the measurement errors and the variance of the measurement error in household consumption.

In the case of a log-linear equation such as that of equation (13), Deaton (1997) notes that elasticity estimates using the log of nonfood expenditures as the sole instrumental variable (IV) are likely to be biased downward, implying that the elasticities estimated by OLS and IV methods provide upper and lower bounds, respectively, for their true value. To address these considerations, the analysis uses an index of household assets, constructed using the principal components method, as an instrumental variable for \( \ln PCE_{ijkt} \). Specifically, in each survey year the index of household assets is estimated using variables for the number of cows and buffaloes, sheep and goats, chickens and ducks, and pigs owned by the household.

In addition, a series of dummy variables summarizes the household residence and its environment, such as whether the roof is concrete or tile, the walls are brick or wood, the floors are tile or cement, the toilet is private or shared, drinking water is accessed through a public network or purchased, and energy for cooking, lighting, and heating is obtained through the public electric or gas utility. Because the consumption variable interacting with the 1999 dummy variable enters into the estimated equation, the instrument used for this interaction term is the asset index based on 1999 data interacting with the 1999 year dummy variable. The results of the first-stage regressions are in appendix B.21 To examine whether the estimated elasticities vary significantly...

21. In all cases, a Hausman-type test (Hausman 1978; Holly 1982) for the absence of measurement error in the consumption variable rejected the null hypothesis.
for the poor, separate analyses are performed for the entire sample and for the sample of households with per capita expenditure below the median.

Table 1 presents the elasticity estimates for SSR obtained from these regressions, fitted separately for the entire sample and for the subset of poorer households. As expected, the negative sign on the point estimates for SSR income elasticity is consistent with the nonparametric observation that the share of calories derived from starchy staples declines with income. Further, the comparison of the elasticity estimates of the IV estimates between urban and rural areas suggests that SSR elasticity is marginally higher in rural Central Java. This implies that the rate at which households, as their incomes rise, switch from cheap sources of calories to food groups that are better sources of micronutrients is higher in rural than in urban areas.

The estimated coefficient for $\beta_{99}$ in the regression is of particular interest because it contains information on how different the elasticities were in 1999 (the crisis year) relative to 1996 (the reference year). The estimates for $\beta_{99}$ in the IV specification reported in table 1 suggest that SSR elasticity in 1999 was practically identical to the elasticity in 1996. The size of the marginal effect is

**Table 1. Elasticities of the Starchy Staple Ratio for Rural and Urban Households in Central Java, 1999**

<table>
<thead>
<tr>
<th>Area and variable</th>
<th>All households</th>
<th>Poorer households*</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Ordinary least squares</td>
<td>Instrumental variable</td>
</tr>
<tr>
<td>Rural lnPCE</td>
<td>$-0.25^{***}$</td>
<td>$-0.24^{***}$</td>
</tr>
<tr>
<td></td>
<td>(0.01)</td>
<td>(0.01)</td>
</tr>
<tr>
<td>Marginal effect</td>
<td>0.01</td>
<td>0.00</td>
</tr>
<tr>
<td></td>
<td>(0.02)</td>
<td>(0.02)</td>
</tr>
<tr>
<td>Urban lnPCE</td>
<td>$-0.26^{***}$</td>
<td>$-0.22^{***}$</td>
</tr>
<tr>
<td></td>
<td>(0.02)</td>
<td>(0.02)</td>
</tr>
<tr>
<td>Marginal effect</td>
<td>$-0.02$</td>
<td>0.00</td>
</tr>
<tr>
<td></td>
<td>(0.02)</td>
<td>(0.04)</td>
</tr>
</tbody>
</table>

*** Significant at the 1 percent level.

**Note:** Numbers in parentheses are robust standard errors, corresponding to the elasticity estimates. Each column represents a separate regression using a wide range of household-level economic and demographic control variables. Instrumental variable estimates were obtained by instrumenting the natural logarithms of per capita expenditure with household-specific asset indexes.

a. Defined as the lower half of the distribution based on consumption per capita.

**Source:** Authors’ analysis based on SUSENAS 1996 and 1999 data.
small and not statistically significant—implying that the income elasticity of SSR was invariant to the different relative prices prevailing in 1999.

Comparing these results with those for the subset sample of the poor in Central Java shows a generally higher elasticity among the poor. This indicates that cash transfers could be more effective in reducing the SSR for the poor relative to the entire sample population. But the fact that the coefficients of the interaction term are not significantly different from zero implies that cash transfers may have been ineffective during the crisis year.

Even though the invariance of SSR elasticity to price increases suggests little substitution toward or away from cereals and tubers, the complex pattern of price changes in Indonesia during the financial crisis of the late 1990s could have induced households to substitute within and among other food groups. To probe deeper into the consequences of such substitutions for the elasticity of micronutrients, the estimation in equation (13) is repeated using the natural logarithm of the consumption of specific nutrients as dependent variables. The elasticity coefficients obtained from these regressions for the entire sample are reported in table 2.

All the OLS estimates of the nutrient elasticities for both rural and urban Central Java are positive and statistically significant for 1996. For example, in rural areas the estimates for the elasticity range from 0.18 for carbohydrates to 0.63 for fat. In urban areas the spread is narrower, ranging from 0.13 for carbohydrates to 0.44 for fats. The IV estimates, on the other hand, while generally statistically significant, appear to be smaller than the corresponding OLS estimates for both urban and rural areas. This suggests that the upward bias from the correlated errors may outweigh the standard downward attenuation bias in the OLS estimates. It is also noteworthy that vitamin C is the only nutrient for which the IV estimates of elasticity are not significant in both urban and rural areas—suggesting that vitamin C consumption may not be responsive to income in Central Java, whether in a crisis year or not.

The estimates for $\beta_{99}$, the coefficient on the interaction between the logarithm of per capita expenditure and the year dummy for 1999, are statistically significant for calories, proteins, fats, carbohydrates, phosphorous, iron, and vitamin B for rural areas and for calories, proteins, fats, calcium, phosphorous, and iron for urban areas. This is evidence of a difference in the elasticity of these nutrients in the crisis year relative to the reference year. Moreover, these differences are quite large, particularly for urban areas. The income elasticities of calories and iron appear to have doubled in urban areas in 1999. The elasticities for proteins, calcium, and phosphorous also nearly doubled. The income elasticity of fats more than doubled in 1999 for urban households.

But there are some nutrients for which income elasticity did not change significantly between 1996 and 1999: calcium and vitamin A in rural areas and carbohydrates, vitamin A, and vitamin B in urban areas. An increase in income elasticity of any specific nutrient in the crisis year—particularly of the magnitudes seen in urban areas—may be considered an indicator of deteriorating
Table 2. Income Elasticities of Nutrients for Rural and Urban Households in Central Java, 1999

<table>
<thead>
<tr>
<th>Area, estimator, and variable</th>
<th>Calories</th>
<th>Proteins</th>
<th>Fats</th>
<th>Carbohydrates</th>
<th>Calcium</th>
<th>Phosphorus</th>
<th>Iron</th>
<th>Vitamin A</th>
<th>Vitamin B</th>
<th>Vitamin C</th>
</tr>
</thead>
<tbody>
<tr>
<td>Rural Ordinary least squares</td>
<td></td>
<td></td>
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<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>InPCE</td>
<td>0.25***</td>
<td>0.31***</td>
<td>0.63***</td>
<td>0.18***</td>
<td>0.37***</td>
<td>0.22***</td>
<td>0.28***</td>
<td>0.27***</td>
<td>0.19***</td>
<td>0.20***</td>
</tr>
<tr>
<td>(0.02)</td>
<td>(0.02)</td>
<td>(0.04)</td>
<td>(0.01)</td>
<td>(0.03)</td>
<td>(0.02)</td>
<td>(0.03)</td>
<td>(0.06)</td>
<td>(0.02)</td>
<td>(0.06)</td>
<td></td>
</tr>
<tr>
<td>Marginal effect (lnPCE*D99)</td>
<td>0.05**</td>
<td>0.09***</td>
<td>0.03</td>
<td>0.07***</td>
<td>0.10***</td>
<td>0.16***</td>
<td>–0.08</td>
<td>0.11***</td>
<td>–0.07</td>
<td></td>
</tr>
<tr>
<td>(0.03)</td>
<td>(0.03)</td>
<td>(0.06)</td>
<td>(0.02)</td>
<td>(0.04)</td>
<td>(0.03)</td>
<td>(0.04)</td>
<td>(0.08)</td>
<td>(0.04)</td>
<td>(0.09)</td>
<td></td>
</tr>
<tr>
<td>Instrumental variable InPCE</td>
<td>0.14***</td>
<td>0.19***</td>
<td>0.50***</td>
<td>0.07***</td>
<td>0.22***</td>
<td>0.10***</td>
<td>0.16***</td>
<td>0.13***</td>
<td>0.07***</td>
<td>–0.03</td>
</tr>
<tr>
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<td>(0.02)</td>
<td>(0.04)</td>
<td>(0.02)</td>
<td>(0.03)</td>
<td>(0.02)</td>
<td>(0.03)</td>
<td>(0.05)</td>
<td>(0.02)</td>
<td>(0.05)</td>
<td></td>
</tr>
<tr>
<td>Marginal effect (lnPCE*D99)</td>
<td>0.07**</td>
<td>0.10**</td>
<td>0.19***</td>
<td>0.06**</td>
<td>0.07</td>
<td>0.14***</td>
<td>–0.15</td>
<td>0.13**</td>
<td>–0.10</td>
<td></td>
</tr>
<tr>
<td>(0.03)</td>
<td>(0.04)</td>
<td>(0.07)</td>
<td>(0.03)</td>
<td>(0.05)</td>
<td>(0.04)</td>
<td>(0.04)</td>
<td>(0.09)</td>
<td>(0.05)</td>
<td>(0.09)</td>
<td></td>
</tr>
<tr>
<td>Urban Ordinary least squares</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>InPCE</td>
<td>0.19***</td>
<td>0.26***</td>
<td>0.44***</td>
<td>0.13***</td>
<td>0.36***</td>
<td>0.22***</td>
<td>0.25***</td>
<td>0.30***</td>
<td>0.19***</td>
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</tr>
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<td>(0.01)</td>
<td>(0.02)</td>
<td>(0.04)</td>
<td>(0.01)</td>
<td>(0.03)</td>
<td>(0.02)</td>
<td>(0.03)</td>
<td>(0.05)</td>
<td>(0.02)</td>
<td>(0.05)</td>
<td></td>
</tr>
<tr>
<td>Marginal effect (lnPCE*D99)</td>
<td>0.06*</td>
<td>0.06</td>
<td>0.23***</td>
<td>0.04</td>
<td>0.11**</td>
<td>0.05</td>
<td>0.11**</td>
<td>0.06</td>
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<td>(0.04)</td>
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<td>(0.08)</td>
<td>(0.05)</td>
<td>(0.08)</td>
<td></td>
</tr>
<tr>
<td>Instrumental variable InPCE</td>
<td>0.12***</td>
<td>0.18***</td>
<td>0.34***</td>
<td>0.07***</td>
<td>0.26***</td>
<td>0.15***</td>
<td>0.17***</td>
<td>0.22***</td>
<td>0.13***</td>
<td>0.09</td>
</tr>
<tr>
<td>(0.02)</td>
<td>(0.02)</td>
<td>(0.04)</td>
<td>(0.02)</td>
<td>(0.03)</td>
<td>(0.02)</td>
<td>(0.03)</td>
<td>(0.06)</td>
<td>(0.03)</td>
<td>(0.06)</td>
<td></td>
</tr>
<tr>
<td>Marginal effect (lnPCE*D99)</td>
<td>0.12**</td>
<td>0.16**</td>
<td>0.38***</td>
<td>0.08</td>
<td>0.23***</td>
<td>0.14**</td>
<td>0.17**</td>
<td>0.07</td>
<td>0.08</td>
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<tr>
<td>(0.05)</td>
<td>(0.07)</td>
<td>(0.10)</td>
<td>(0.05)</td>
<td>(0.08)</td>
<td>(0.07)</td>
<td>(0.07)</td>
<td>(0.12)</td>
<td>(0.08)</td>
<td>(0.11)</td>
<td></td>
</tr>
</tbody>
</table>

*** Significant at the 1 percent level; ** significant at the 5 percent level; * significant at the 10 percent level.

Note: Numbers in parentheses are robust standard errors, corresponding to the elasticity estimates. Each column represents a separate regression using a wide range of household-level economic and demographic control variables. All specifications control for cluster fixed effects and year dummy variables. Instrumental variable estimates were obtained by instrumenting the natural logarithms of per capita expenditure with household-specific asset indexes.

Source: Authors’ analysis based on SUSENAS 1996 and 1999 data.
nutritional quality and availability. Consider the income elasticity of fats, which have the highest elasticity of all nutrients for both rural and urban consumers. Among households in Central Java, this reflects the “luxury good” status of foods such as meat and fish. The price increases in 1999 also appear to have severely curtailed households’ ability to obtain other basic nutrients, such as protein, calcium, phosphorous, and iron, as well as overall calories.

Table 3 shows elasticity estimates for a restricted sample of the poor (per capita expenditure below the median), which are uniformly higher than for the entire sample population. The IV estimates reveal four broad types of nutrients: those for which income elasticities are statistically indistinguishable from zero in 1996 and 1999, those for which elasticities are statistically indistinguishable from zero in 1996 but significant in 1999, those for which elasticities are statistically different from zero but not statistically different between 1996 and 1999, and those for which elasticities are statistically different from zero and statistically different between the two years.

Table 4 categorizes nutrients based on the IV results for Central Java. It shows that, unlike for the entire sample, the elasticity estimates for poorer households suggest considerable heterogeneity in the elasticities of nutrients. For the urban poor in Central Java, for example, elasticity in the normal year (1996) is significantly different from zero only for fats. But in the crisis year (1999), elasticity is significantly larger for fats as well as for calories, proteins, carbohydrates, phosphorous, and vitamin B. Elasticities for calcium, iron, and vitamins A and C are never significant, indicating that cash transfers may not be effective vehicles for protecting these nutrients in poorer households.22

Robustness Check: Is It Prices or Other Factors?

The empirical specification of equation (13) relies on cross-sectional variation in relative prices to identify income elasticities. Pooling the cross-sectional data from the two survey years allowed the construction of a basic test for whether the estimated elasticities were different in 1996 and 1999. Although a dummy variable for year is included in the specification, which would presumably absorb everything else that changed between the two survey years, there could be lingering questions about whether the findings are due to changes in food prices specifically or perhaps driven by other, unrelated effects of the economic crisis. If food prices could be observed at the household level, then the issue could be tested directly. But prices are not observed at the household level. Rather, unit values are observed, and these are known to be contaminated by variations in quality. To confirm that changes in food prices are indeed driving these results,

22. A simple calculation based on IV estimates for all households in urban Central Java shows that a cash transfer of 25 percent of income would increase the consumption of nutrients considered in this article by from 6–18 percent in the 1999 crisis year—twice the impact range of 3–9 percent for the same set of nutrients in the 1996 noncrisis year.
### Table 3. Income Elasticities of Nutrients for Poorer Households in Rural and Urban Central Java, 1999

<table>
<thead>
<tr>
<th>Area, estimator, and variable</th>
<th>Calories</th>
<th>Proteins</th>
<th>Fats</th>
<th>Carbohydrates</th>
<th>Calcium</th>
<th>Phosphorus</th>
<th>Iron</th>
<th>Vitamin A</th>
<th>Vitamin B</th>
<th>Vitamin C</th>
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<tbody>
<tr>
<td><strong>Rural</strong></td>
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<td></td>
<td></td>
<td></td>
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</tr>
<tr>
<td>lnPCE</td>
<td>0.29***</td>
<td>0.36***</td>
<td>0.74***</td>
<td>0.22***</td>
<td>0.45***</td>
<td>0.24***</td>
<td>0.35***</td>
<td>0.28***</td>
<td>0.19***</td>
<td>0.23**</td>
</tr>
<tr>
<td>(0.02)</td>
<td>(0.02)</td>
<td>(0.06)</td>
<td>(0.02)</td>
<td>(0.04)</td>
<td>(0.03)</td>
<td>(0.04)</td>
<td>(0.09)</td>
<td>(0.03)</td>
<td>(0.03)</td>
<td>(0.10)</td>
</tr>
<tr>
<td>Marginal effect (lnPCE*D99)</td>
<td>0.12***</td>
<td>0.14***</td>
<td>0.08</td>
<td>0.14***</td>
<td>0.11*</td>
<td>0.16***</td>
<td>0.23***</td>
<td>–0.04</td>
<td>0.22***</td>
<td>–0.15</td>
</tr>
<tr>
<td>(0.03)</td>
<td>(0.04)</td>
<td>(0.09)</td>
<td>(0.03)</td>
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<td>(0.13)</td>
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<td>Instrumental variable</td>
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</tr>
<tr>
<td>lnPCE</td>
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<td>0.41***</td>
<td>–0.03</td>
<td>0.11*</td>
<td>–0.14</td>
<td>0.05</td>
<td>–0.05</td>
<td>–0.12**</td>
<td>0.37***</td>
</tr>
<tr>
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<td>(0.04)</td>
<td>(0.07)</td>
<td>(0.04)</td>
<td>(0.06)</td>
<td>(0.08)</td>
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<td>(0.11)</td>
<td>(0.05)</td>
<td>(0.05)</td>
<td>(0.12)</td>
</tr>
<tr>
<td>Marginal effect (lnPCE*D99)</td>
<td>0.04</td>
<td>0.06</td>
<td>0.36**</td>
<td>0.01</td>
<td>0.17</td>
<td>0.03</td>
<td>0.19*</td>
<td>–0.05</td>
<td>0.18</td>
<td>–0.16</td>
</tr>
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<td>(0.09)</td>
<td>(0.16)</td>
<td>(0.08)</td>
<td>(0.11)</td>
<td>(0.17)</td>
<td>(0.10)</td>
<td>(0.24)</td>
<td>(0.13)</td>
<td>(0.13)</td>
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<tr>
<td><strong>Urban</strong></td>
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<tr>
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<tr>
<td>lnPCE</td>
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<td>0.31***</td>
<td>0.79***</td>
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<td>0.49***</td>
<td>0.24***</td>
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<td>0.41***</td>
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<td>(0.03)</td>
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<td>(0.03)</td>
<td>(0.03)</td>
<td>(0.05)</td>
<td>(0.05)</td>
<td>(0.04)</td>
<td>(0.13)</td>
</tr>
<tr>
<td>Marginal effect (lnPCE*D99)</td>
<td>0.19***</td>
<td>0.14***</td>
<td>0.23**</td>
<td>0.20***</td>
<td>0.10</td>
<td>0.17***</td>
<td>0.16**</td>
<td>–0.04</td>
<td>0.21***</td>
<td>–0.10</td>
</tr>
<tr>
<td>(0.05)</td>
<td>(0.05)</td>
<td>(0.10)</td>
<td>(0.05)</td>
<td>(0.09)</td>
<td>(0.05)</td>
<td>(0.07)</td>
<td>(0.17)</td>
<td>(0.06)</td>
<td>(0.06)</td>
<td>(0.19)</td>
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<td>Instrumental variable</td>
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<td></td>
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<td></td>
<td></td>
<td></td>
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</tr>
<tr>
<td>lnPCE</td>
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<td>0.05</td>
<td>0.47***</td>
<td>–0.11</td>
<td>0.22</td>
<td>–0.00</td>
<td>0.06</td>
<td>0.06</td>
<td>–0.08</td>
<td>–0.17</td>
</tr>
<tr>
<td>(0.07)</td>
<td>(0.08)</td>
<td>(0.17)</td>
<td>(0.08)</td>
<td>(0.15)</td>
<td>(0.08)</td>
<td>(0.13)</td>
<td>(0.31)</td>
<td>(0.10)</td>
<td>(0.33)</td>
<td></td>
</tr>
<tr>
<td>Marginal effect (lnPCE*D99)</td>
<td>0.38***</td>
<td>0.36**</td>
<td>0.76***</td>
<td>0.35**</td>
<td>0.15</td>
<td>0.35**</td>
<td>0.18</td>
<td>0.18</td>
<td>0.35*</td>
<td>–0.05</td>
</tr>
<tr>
<td>(0.13)</td>
<td>(0.15)</td>
<td>(0.27)</td>
<td>(0.15)</td>
<td>(0.24)</td>
<td>(0.15)</td>
<td>(0.21)</td>
<td>(0.51)</td>
<td>(0.19)</td>
<td>(0.32)</td>
<td></td>
</tr>
</tbody>
</table>

*** Significant at the 1 percent level; ** significant at the 5 percent level; * significant at the 10 percent level.

Note: Numbers in parentheses are robust standard errors, corresponding to the elasticity estimates. Each column represents a separate regression using a wide range of household-level economic and demographic control variables. All specifications control for cluster fixed effects and year dummies. Instrumental variable estimates were obtained by instrumenting natural logarithms of per capita expenditure with household-specific asset indexes. Poorer household are defined as the lower half of the distribution based on per capita expenditure.

Source: Authors’ analysis based on SUSENAS 1996 and 1999 data.
regressions of the following form are analyzed for each nutrient:

\[
\ln Y_j = \beta_0 + \beta_1 \ln PCE_j + \sum_i \beta_i \ln \left( \frac{P_i}{P_{CT}} \right) + \sum_i \gamma_i \ln \left( \frac{P_i}{P_{CT}} \right) \times \ln PCE_j \\
+ \delta' X_j + \epsilon_j
\]

(14)

where \( i \) indexes four food groups: meat and fish, fruits and vegetables, eggs and milk, and others, \( P_{CT} \) is the average of the unit values of cereals and tubers at the village (desa) level, and \( P_i \) is the average unit value of each of the four food groups such that the ratios in equation (14) represent the price of each of the four food groups relative to the price of cereals and tubers. As before, \( X_j \) represents the set of controls used in previous regressions. Using village-level averages of unit values avoids the problem of quality differences in household-level purchases of food.\(^{23}\)

These regressions are estimated for 1996 with the primary goal of assessing whether or not the coefficients on the interaction terms (the \( \gamma_i \)'s) are significantly different from zero. These estimations are made for each of the micronutrients and separately for urban and rural central Java. These estimations reveal that at least one interaction term is always statistically different from zero. In addition, an F-test of the joint significance of these interactions rejects the null hypothesis in almost all instances (see Appendix S4, a supplemental appendix available at http://wber.oxfordjournals.org). This makes an explicit connection between the estimated income elasticity of any one micronutrient and the

\(^{23}\) Deaton (1988) notes that using cluster means of unit values in regressions of this form is essentially the same as using cluster dummy variables to instrument for individual unit values in a regression of shares and prices, where prices are expressed as unit values.
prevailing relative prices of food in the economy, corroborating the finding that elasticities estimated under one set of relative prices do not necessarily remain valid when relative prices change dramatically. In particular, this also provides additional confidence in the attribution of differences in the income elasticities of micronutrients between the normal year and the crisis year to food prices.

III. IMPLICATIONS

There is considerable heterogeneity in the income elasticity of demand for nutrients over time based on analysis of household data from the 1996 and 1999 consumption modules of the SUSENAS in Indonesia. A comparison of OLS and IV estimates of the demand for nutrients suggests that OLS estimates are likely to be misleading due to bias from correlated errors in consumption and nutrient content. In particular, the finding that IV estimates are generally lower than OLS estimates suggests that the upward bias due to correlated measurement errors in nutrient intake and household-level consumption may outweigh the possible attenuation bias.

The analysis also shows that for most nutrients, including micronutrients such as phosphorous, iron, and calcium, income elasticity estimates are significantly higher in a crisis year. Moreover, the magnitude of the increase appears to be larger in urban than in rural areas. On the other hand, for some micronutrients, such as vitamin C, the income elasticity estimates obtained from the IV specification are statistically indistinguishable from zero. This suggests that income may have limited leverage to increase or protect the consumption of vitamin C, whether in a crisis year or not. The separate analysis of nutrient elasticity for poor households reinforces the message that, some nutrients could be effectively protected using cash transfers during crisis years (such as iron and fats in rural Central Java) while others are unlikely to be responsive to income supplements (such as vitamin A in rural Central Java).

These results have two specific policy implications. First, given the significant increases in the income elasticities of both micronutrients and macronutrients during economic crises suggests that cash transfer programs can help households protect their consumption of essential nutrients, with important differences between the urban and rural poor. To the extent that delivery infrastructure is already in place exists and leakage is low, cash transfers are widely accepted as the quickest and cheapest intervention to scale up to reach households most likely to be adversely affected. This research shows that they can also be more effective in protecting the consumption of some key nutrients during economic crises than in normal economic conditions.

Second, a complete reliance on cash transfers may be insufficient if the policy goal of policy in response to economic crises is to protect all important nutrients. For example, consumption of vitamin C, an important micronutrient, was unresponsive to income in both rural and urban Central Java. This suggests that targeted micronutrient supplementation programs, designed with careful attention to
differences between the urban and rural poor, might have to accompany cash transfers to ensure that key micronutrients are not sacrificed during crises. Future research could be directed at understanding and identifying specific nutrients that households are likely to sacrifice during a crisis in different settings.

**Appendix A. Data Description**

Data from the consumption module of SUSENAS, collected every three years, included 216 food items in 1996 and 214 food items in 1999. The seven-day food intake survey makes a very good effort to capture the total value of the food consumed by households. In both years, households were asked to report the quantity and value of food purchased, given to them as gifts, or consumed out of their own production during the previous week. Items are valued by local interviewers using the prevailing market prices in the villages where households reside.

The micronutrient content of each food item is calculated using conversion factors published by the Nutrition Directorate in the Ministry of Health of Indonesia (Direktorat Gizi, Departemen Kesehatan 1988). That publication contains the micronutrient content of a comprehensive list of foods; each was matched with one or more of the food items captured in the SUSENAS consumption module. Because both the quantities and calories of each food item are available in the SUSENAS dataset, either may be used to derive the micronutrient content. Additional investigation determined that it is preferable to rely on the calorie data rather than the quantity data. In 1996, for example, the quantity of various food items was recorded in kilograms rather than in grams as the questionnaire specified. In addition, for a number of food items, quantity was coded in pieces, such as number of eggs, rather than in weight, but calories were provided per unit of weight. Similar problems were noted with the coding of quantities in 1999.

Because of these issues, the analysis in this article used the calorie information provided by the BPS to derive a more reliable measure of the quantity of each food item and micronutrient consumed. First, the standard calories-to-quantity conversion formula (also applied by the Central Statistical Agency) was used to derive a new quantity for each food item. Second, the quantity-to-micronutrient formula obtained from the Ministry of Health was applied to derive the quantity of micronutrients for each food item. This approach implicitly assumes that the calorie data are more reliable than the

24. The difference arises from the fact that “high quality” and “imported” rice were treated as separate food items in the cereals category in 1996, but not in 1999.
25. Van de Walle (1988) provides a guide to the SUSENAS consumption module that is still very useful despite some changes in the questionnaire.
27. In 1996 this was the case for food items with codes 75, 76, 81, 82, 84, 157, 167, and 203.
original quantity data—a reasonable assumption because the quantity data must have been processed in some way to apply the standard conversion factors to calculate calories.

The value of food consumption is the sum of spending on grains, meat and fish, eggs and milk, vegetables, pulses, fruits, seasonings, fats and oils, soft drinks, prepared foods and other food items, and alcohol.\textsuperscript{28} The reference period for consumption of these items is the seven days preceding the day of the survey interview. Weekly consumption was transformed into monthly consumption transformed into monthly consumption by multiplying by \((30/7)\) For nonfood expenditures the survey collects two measures, one for the prior month and one for the previous 12 months. To avoid exclusion errors, average expenditures per month were calculated from the reported expenditures of the last 12 months. Expenditures on nonfood items include tobacco, housing, clothing, health and personal care, education and recreation, transportation and communication, taxes and insurance, and ceremonial expenses. Expenditures on durables such as household furniture, electric appliances, and audiovisual equipment are excluded for aggregate household consumption. A household’s income is measured by monthly per capita consumption, denoted by and calculated by dividing the monthly total of food and nonfood consumption in survey period \(t\) by the size of the household in the period.\textsuperscript{29}

\section*{Appendix B. First-stage Regressions}

\begin{table}[h]
\centering
\begin{tabular}{|l|l|l|l|l|l|}
\hline
\textbf{Variable} & \textbf{(1)} & \textbf{(2)} & \textbf{(1)} & \textbf{(2)} \\
\hline
Log of household asset index & 0.24*** & -0.02*** & Household head self-employed with permanent assistance & 0.20*** & 0.01 \\
 & (0.01) & (0.00) & & (0.04) & (0.02) \\
Log of household asset index*D99 & -0.12*** & 0.14*** & Household head working without pay & 0.00 & 0.00 \\
 & (0.01) & (0.01) & & (0.02) & (0.01) \\
Number of boys under age 5 & 0.03 & 0.04 & Household head literate & 0.00 & 0.04* \\
 & (0.08) & (0.03) & & (0.07) & (0.02) \\
\hline
\end{tabular}
\end{table}

(Continued)

\textsuperscript{28} Unlike SUSENAS, this article does not include tobacco expenditures in the food consumption total.

\textsuperscript{29} It is implicitly assumed that there are no economies of scale at the household level. For comparing income elasticity over time, this assumption is not overly limiting. In any case, the regression analysis controls for the gender and age composition of families in each survey year.
<table>
<thead>
<tr>
<th>Variable</th>
<th>Column 1</th>
<th>Column 2</th>
<th>Variable</th>
<th>Column 1</th>
<th>Column 2</th>
</tr>
</thead>
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<tr>
<td>Number of girls under age 5</td>
<td>-0.11</td>
<td>0.02</td>
<td>Spouse has no education</td>
<td>-0.04</td>
<td>0.01</td>
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<td>(0.08)</td>
<td>(0.03)</td>
<td></td>
<td>(0.03)</td>
<td>(0.01)</td>
</tr>
<tr>
<td>Number of boys ages 6–12</td>
<td>0.07</td>
<td>0.02</td>
<td>Spouse has not completed primary school</td>
<td>-0.09***</td>
<td>0.03***</td>
</tr>
<tr>
<td></td>
<td>(0.08)</td>
<td>(0.03)</td>
<td></td>
<td>(0.03)</td>
<td>(0.01)</td>
</tr>
<tr>
<td>Number of girls ages 6–12</td>
<td>0.06</td>
<td>0.02</td>
<td>Spouse has completed primary school</td>
<td>-0.10***</td>
<td>0.02**</td>
</tr>
<tr>
<td></td>
<td>(0.08)</td>
<td>(0.03)</td>
<td></td>
<td>(0.03)</td>
<td>(0.01)</td>
</tr>
<tr>
<td>Number of boys ages 13–19</td>
<td>0.26***</td>
<td>0.01</td>
<td>Spouse has completed junior or senior high school</td>
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<td>0.01</td>
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<tr>
<td></td>
<td>(0.08)</td>
<td>(0.03)</td>
<td></td>
<td>(0.03)</td>
<td>(0.01)</td>
</tr>
<tr>
<td>Number of girls ages 13–19</td>
<td>0.27***</td>
<td>0.03</td>
<td>Spouse self-employed without assistance</td>
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<td>0.01</td>
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<tr>
<td></td>
<td>(0.08)</td>
<td>(0.03)</td>
<td></td>
<td>(0.02)</td>
<td>(0.01)</td>
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<tr>
<td>Number of men ages 20–54</td>
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<td>Spouse self-employed with nonpermanent assistance</td>
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<td>(0.07)</td>
<td>(0.02)</td>
<td></td>
<td>(0.03)</td>
<td>(0.01)</td>
</tr>
<tr>
<td>Number of women ages 20–54</td>
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<td>0.01</td>
<td>Spouse self-employed with permanent assistance</td>
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<td>-0.09**</td>
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<td>(0.06)</td>
<td>(0.02)</td>
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<td>(0.06)</td>
<td>(0.05)</td>
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<tr>
<td>Household head has no education</td>
<td>-0.20***</td>
<td>0.05**</td>
<td>Spouse working without pay</td>
<td>0.08**</td>
<td>0.03*</td>
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<td>(0.02)</td>
<td></td>
<td>(0.04)</td>
<td>(0.01)</td>
</tr>
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<td>Household head has not completed primary school</td>
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<td>Spouse with wage employment</td>
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<td>0.00</td>
</tr>
<tr>
<td></td>
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<td>(0.01)</td>
<td></td>
<td>(0.02)</td>
<td>(0.01)</td>
</tr>
<tr>
<td>Household head has completed primary school</td>
<td>-0.09***</td>
<td>0.01</td>
<td>Spouse literate</td>
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<td>-0.01</td>
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<td>(0.01)</td>
<td></td>
<td>(0.03)</td>
<td>(0.01)</td>
</tr>
<tr>
<td>Household head has completed junior or senior high school</td>
<td>-0.06**</td>
<td>-0.01</td>
<td>Household size</td>
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(Continued)
Table B-1. Continued

<table>
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<th>Household head self-employed without assistance</th>
<th>-0.01</th>
<th>0.01</th>
<th>D99 (=1 if year is 1999)</th>
<th>0.58***</th>
<th>12.52***</th>
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<td>(0.14)</td>
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<table>
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<tr>
<th>Household head self-employed with nonpermanent assistance</th>
<th>0.04*</th>
<th>0.01</th>
<th>R-squared</th>
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<th>0.999</th>
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<td>(0.01)</td>
<td>F-statistic</td>
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<td>24566</td>
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</table>

*** Significant at the 1 percent level; ** significant at the 5 percent level; * significant at the 10 percent level.

Note: These regressions also include the control variables that interact with the 1999 year dummy variable.

Source: Authors’ analysis based on equation (13).

References


