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Can We Rely on Cash Transfers to Protect Dietary Diversity during Food Crises?

Estimates from Indonesia

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Abstract

The 2008 "food price crisis" and more recent spikes in food prices have led to a greater focus on policies and programs to cushion their impact on poverty and malnutrition. Estimating the income elasticity of micro-nutrients and assessing how they change during such crises is an important part of the policy debate as it affects the effectiveness of cash transfer and nutritional supplementation programs. This paper assesses these issues using data from two cross-sectional household surveys in Indonesia carried out before and soon after the 1997/98 economic crisis, which led to a sharp increase in food prices. First, the authors examine how the income elasticity of the starchy staple ratio differs between the two survey rounds using non-parametric as well as regression methods. Second, they provide updated estimates of the income elasticity for important nutrients in Indonesia. The analysis finds that (i) summary measures such as the income elasticity of the starchy staple ratio may not change during crises but this masks important differences across specific nutrients; (ii) methods matter—the ordinary least squares estimates for the income elasticity of micro-nutrients are likely to be misleading due to measurement error bias; (iii) controlling for measurement error, the income elasticity of some key micro-nutrients, such as iron, calcium, and vitamin B1, is significantly higher in the crisis year compared with a normal year; and (iv) the income elasticity for certain micro-nutrients—vitamin C in this case—remains close to zero. These results suggest that cash transfer programs may be even more effective during crises to protect the consumption of many essential micro-nutrients compared with non-crisis periods but in order to ensure that all micro-nutrients are consumed, specific nutritional supplementation programs are also likely to be required.

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Can we rely on cash transfers to protect dietary diversity during food crises? Estimates from Indonesia

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1. Introduction

International food prices increased sharply in 2008 and in real terms, reached levels not seen since the early 1970s. The FAO Food Price Index grew by 73% between September 2006 and mid 2008 and was driven by unprecedented increases across all food categories. During the same period, meat prices increased by 25%, dairy by 91% oils and fat by 149% and cereal grains increased by 123%[2]. Following a downward trend in global food prices in 2009, food prices rose again in late 2010 and in December 2010 are close to the 2008 peak. Hence food price volatility, and in particular sharp spikes in food prices, needs to be monitored closely for their impact on the poor. A review of the literature on the impact 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 possible adverse effects have not only heightened concerns about food security and the deepening of malnutrition in parts of the developing world, but have also sparked a renewed interest in the design of the most effective policy response. From the point of view of the household, such price increases have two major consequences. Firstly, they result in a decrease in the purchasing power of household income especially among poorer households that spend a larger share of their income on food. Secondly, they result in a relative price effect that induces households to substitute away from the more expensive foods[3]. Government intervention has almost always been motivated by the need to compensate poor households for their lost purchasing power. These interventions – or "social safety net" programs as they are commonly known – are aimed at smoothing consumption and protecting the caloric availability within households to prevent sudden increases in poverty and hunger. A review of the safety net programs used for the 2008 food price crisis show that they typically take the form of income support of various forms – cash transfers and price subsidies – as well as in-kind transfers of staple foods or supplementary feeding programs (Wodon and Zaman 2010).

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[2] Food Outlook, November 2008. Available at: http://www.fao.org/docrep/011/ai474e16htm

[3] The purchasing power will actually improve for net sellers of agricultural commodities whose prices increase. Also, for households that are close to subsistence and are already consuming the cheapest sources of calories, the substitution possibilities are going to be much more limited.

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

Although there is plenty of evidence on the size of the calorie income elasticity (e.g. Strauss and Thomas, 1995, Subramanian and Deaton, 1996, Skoufias, 2003), empirical evidence on the micro-nutrient income elasticity is lacking. Moreover the scarce evidence that exists suggests substantial differences in micro-nutrient-income elasticities (e.g. Behrman and Deolalikar, 1987, Bouis, 1991). Focusing just on Indonesia, for example, Pitt and Rosenzweig (1985), using data from farm households report very low nutrient-income elasticities (below 0.03) for a set of nutrients that includes calories, protein, fat, carbohydrates, calcium, phosphorus, iron, vitamin A, and vitamin C. Another study using data from rural and urban areas reports much higher nutrient income elasticities (for example, from 0.70 to 1.20 for the lower 40 percent of the population by expenditure on Java (Chernichovsky and Meesook, 1984). Similarly diverse estimates are reported for other countries[4].

In addition, most of the empirical evidence to date sheds light on whether the price sensitivity of the demand for food and nutrients varies with the level of income (Timmer and Alderman, 1979; Timmer, 1981; Pitt, 1983) or whether the income elasticity of calories varies with the level of income (Behrman and Deolalikar, 1987; Ravallion, 1990; Strauss and Thomas, 1995; Subramanian and Deaton, 1996). Yet, we were unable to find evidence that could be

[4] Behrman and Deolalikar (1987), for example, using data from ICRISAT villages, report income elasticity estimates of 0.06 to 0.19 for protein (depending on whether level estimates or differences over time are used), 0.30 or -0.22 for calcium, -0.11 to 0.30 for iron, 0.19 to 2.01 for carotene, -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 ascorbic acid. The Nicaraguan study (Behrman and Wolfe, 1987) reports significant income elasticity estimates in the range of 0.04 to 0.11 for calories, protein, iron and vitamin A (with statistically significant but quantitatively small non-linearities). The Philippine study (Bouis, 1991) reports an iron-income elasticity of 0.44, a calorie income elasticity of 0.16 and insignificant income elasticities for vitamin A and vitamin C.

related to the question of whether the nutrient-income elasticity varies significantly depending on the relative prices faced by households, such as those experienced during food price spikes.

During the financial crisis in Indonesia, the value of the rupiah depreciated dramatically in 1998. The rupiah fell from around 2,400 per US$ in June 1997 to just under Rp 15,000 per US$ in June 1998, finally settling down to Rp 8,000 or 9,000 per US$ by December 1998. These fluctuations in the exchange rate led to large increases in the price of tradable commodities in domestic markets. The consumer price index increased by 107 percent between February 1996 and February 1999. During the same period, the food price index rose by 188 percent. In addition, a number of existing subsidies on consumer goods such as rice, oil and fuel were removed in 1998. It is thus questionable whether estimates of the income elasticity of nutrients obtained from a sample of households observed during pre-crisis years can provide any guidance on how caloric and micro-nutrient availability may respond to additional income (ceteris paribus) during a period with a different set of relative prices. From a policy perspective the sensitivity of the nutrient income elasticity to the relative prices in the economy implies that policies aimed at increasing household income such as employment and cash transfer programs may be less effective in protecting nutritional outcomes depending on the economic conditions prevailing at the time of their implementation.

In this paper we use the shock to food prices in Indonesia to assess various aspects of the relationship between nutrient consumption and prices. We conduct the analysis at two levels. First, using the Starchy Staple Ratio (SSR) as the summary measure of household nutritional welfare, we assess the impact of the dramatic change in food prices on household dietary composition. SSR is defined as the share of caloric availability derived from starchy staple foods (cereals and tubers) and according to Bennett's Law this ratio is inversely related to the importance of inexpensive starches relative to higher quality, more expensive, and micro-nutrient-rich foods (such as meat and fish, fruits and vegetables). Specifically, we examine how the income elasticity of the SSR differs in the two survey rounds characterized by very different relative prices between cereals and other major food groups. We conduct the analysis separately for urban and rural areas in the province of Central Java (one of the poorest provinces in Indonesia) in 1996 and 1999 and report results using non-parametric as well as regression methods.

We supplement this analysis by providing updated estimates of the income elasticity for important nutrients in Indonesia, such as calories, protein, carbohydrate, fat, iron, calcium, phosphorous, and vitamins A (carotene), B1 (thiamin), and C (ascorbic acid). At times of crises cash transfers may be the fastest and least costly method of reaching the households most likely to be adversely affected, if the delivery infrastructure exists and has low levels of leakage. Reliable elasticity estimates can help policy makers determine *ex-ante* whether cash transfer program can be at all effective at increasing nutrient availability among poor households or whether alternative interventions may be necessary. Therefore, particular emphasis is placed on the sensitivity of the elasticity estimates to biases due to measurement error in consumption and the nutrient availability at the household level.

We also provide a test for whether the income elasticity for nutrients varies with the economic conditions faced by households. Ceteris paribus, changes in the relative prices of the staple food items may plausibly give rise to rather unexpected responses to how the demand for nutrients may respond to a cash transfer. For example, in situations where the level of total caloric availability is already low, if the relative price of the staple increases during a crisis, households receiving a cash transfer may choose to spend more of their additional income on that same staple as long as it continues to be the cheapest source of calories and energy (Behrman, 1988; Behrman and Deolalikar, 1989)[5]. Jensen and Miller (2008) have extended this analysis and find evidence supporting Giffen good property of basic staple commodities such as rice in China.

The rest of the paper is structured as follows. In section two we describe the data used for the analysis, the way some key variables were constructed and present some background information on the changes in prices and nutrient availability between 1996 and 1999 in Indonesia. In section three we discuss our empirical strategy and the results of estimation using both non-parametric and regression methods. In section four, we summarize our findings and conclude.

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

2. Background and Data

Our analysis is based on the detailed consumption module of the National Socio-Economic Survey (SUSENAS) collected every three years by the Central Statistical Agency (BPS) of the Government of Indonesia. The consumption module is nationally representative of urban and rural areas within each of the 27 provinces.[6] The 1996 round surveys 60,678 households and the 1999 round 62,217 households. Besides the detailed nature of the survey, one of the main advantages obtained by the comparison of the income elasticity of calories in these two years is the opportunity to examine economic behavior in the context of dramatically different relative price regimes. In February 1999, the month in which the 1999 SUSENAS was conducted, the inflation rate in Indonesia had reached its peak since the start of the financial crisis in late 1997 and its intensification in mid 1998. The additional benefit is that the same questionnaire was applied at the same point in time in each survey year. In this manner the possible influence of seasonal factors in the caloric income relationship as emphasized by Behrman, Foster and Rosenzweig (1997) can be controlled[7].

The consumption module includes 216 food items in 1996 and 214 food items in 1999[8]. The survey is a 7-day food intake survey and makes a very good effort at getting to the total value of the food consumed by households. In each of these years households are asked explicitly to recall the quantity and value of each of these food items purchased from the market during the last week, or given to them as gifts or consumed out of own production[9]. The latter quantities are valued by local interviewers using the prevailing market prices in the villages where households reside.

The micro-nutrient 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, RI, 1988). This publication contains the micro-nutrient content per 100

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[6] We also use some variables from the larger SUSENAS (core survey) containing observations for approximately 205, 000 households.

[7] The fasting month and the Idul Fitri-Lebaran holiday following it is a moving holiday and in 1999 it fell in late January of 1999. We were informed by BPS officials that the survey were conducted two weeks after the Lebaran holiday and as a result the value of household food consumption has little chance of appearing unusually high due to the fasting holiday.

[8] The difference of two items arises from the fact that high quality and imported rice in 1996 were treated as separate food items in the cereals category but not in 1999.

[9] Van de Walle (1988) provides a guide to the SUSENAS consumption module that is still very useful in spite of some changes in the questionnaire.

grams of a long list of food items. Each food item in this publication was matched with one or more of the 200 food items in the consumption module of the SUSENAS. Given that both the quantity and calories of each food item are made available with the SUSENAS data set one has the option to use either the quantities or the calories to derive the micro-nutrient content of each food item. Upon closer investigation of the quantities and calories provided for each food item it was determined that it is preferable to rely on the caloric data rather than the raw quantity data. In 1996, for example, the quantity of a various food items was recorded in kilograms (kg) rather than in grams (gr) as indicated by the questionnaire[10]. 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[11]. Similar problems were noted with the coding of quantities of food items in 1999.

Hence rather than using the raw quantity variable collected by the consumption module in each year we used the caloric information provided by BPS for each food item in each year in order to derive a more reliable measure of the quantity of each food item consumed. Firstly, the standard quantity to calories conversion formula also applied by the BPS was used to derive the "new" quantity of each food item. Secondly, the quantity to micro-nutrient formula obtained by the Ministry of Health was applied to derive the quantity of micro-nutrients available for each food item. Of course this approach implicitly assumes that the caloric data provided by the BPS are more reliable than the original quantity data, an assumption that may be quite reasonable since BPS data processors must have processed the quantity data in one way or another in order to apply the standard calorie conversion factors and generate calories.

The value of food consumption is the sum of expenditures on grains, meat, fish, eggs and milk, vegetables, pulses, fruits, seasonings, fats and oils, soft drinks, prepared food and other food items, and alcohol[12]. The reference period for consumption of these items is the week preceding the day of the interview. Weekly consumption was transformed into monthly consumption by multiplying by (30/7).

For non-food expenditures the survey collects two measures, each for a different reference period: last month and last 12 months. In order to avoid exclusion errors we used the

[10] In 1996 this was the case for food items with codes 45-52, 95, 102, 110-114, 126-127, 158-166, 171-180, 184-185, 187-194, 208, 212 and 215.
[11] In 1996, this is the case for food items with codes 75, 76, 81, 82, 84, 157, 167 and 203.
[12] In contrast to BPS, we do not include tobacco expenditures in the food consumption total.

average expenditures per month calculated from the reported expenditures based on the reference period of the last 12 months. Expenditures on non-food items include expenditures on tobacco, housing, clothing, health and personal care, education and recreation, transportation and communication, taxes and insurance and other ceremonial expenses. Expenditures on durables such as household furniture, electric appliances, and audio-visual equipment, are excluded for the aggregate of household consumption. The income of a household is measured by the value of monthly per capita consumption, denoted by $PCE(t)$ is constructed by dividing the monthly value of total food and plus non-food consumption in survey period t by the size of the household in each period[13].

In order to make any meaningful comparison across two cross-sectional surveys that are three years apart it is essential to express the nominal income of households in 1999 in terms of 1996 rupiah. A critical point for the construction of "real" income in 1999 is the fact that changes in food prices impact on households differently depending on the share of their budget they spend on food. Typically, poorer households spend a much higher fraction of the income on food (closer to 60% for poor rural households in Indonesia), while this share diminishes down to 40 percent for households at the top of the expenditure scale in urban areas.

The availability of the value and quantity for each of the food items in the SUSENAS consumption modules allows calculation of unit values at the household level. Given the data available, we constructed a deflator combining the unit values calculated from the consumption module and the province-specific prices reported for non-food items by the BPS[14]. First, given that for non-food items only expenditures are collected, we constructed a deflator for non-food items using the mean shares of major groups of non-food items in the February 1999 survey as weights and the province-specific price indices for these groups[15]. Second, we constructed a household-specific food deflator from a weighted average of the fifty-two food items used in the calculation of the poverty line in Indonesia. Specifically, the household-specific food deflator is calculated using the formula

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[13] Thus it is implicitly assumed that there are no economies of scale at the household level. For the present purpose of comparing income elasticity over time, this assumption is not overly limiting. In any case, the regression analysis below controls for the gender and age composition of families in each survey year.

[14] More details regarding the construction of the price index can be found in Skoufias (2003). Suryahadi et al. (2000) and Levinson et al. (1999) adopt a similar approach in constructing household specific price indices for Indonesia.

[15] It should be noted that the province-specific price indices for food and non-food groups reported by BPS are based solely on urban prices for 27 cities in 1996 and 44 cities in 1999.

$$P_F^h(99) = \left[\sum_{i=1}^{52}\left(S_i^h(99)\left(\frac{P_i\,(R,96)}{P_i\,(R,99)}\right)\right)\right]^{-1} \qquad (1)$$

which is the standard formula for calculating a Paasche price index (see Deaton and Zaidi, 2002). The letter S denotes the share of food item i of the total amount expended on the 52 food items and the superscript h indicates that this share varies from household to 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 within each of the urban and rural areas of each of the 26 provinces and the Jakarta metropolitan area (a total of 53 regions), so as to minimize the influence of measurement errors and differences in the quality of food consumed by wealthier households (Deaton, 1988). Having a price deflator for food and non-food, the price deflator for household h in 1999, $P^h(99)$, can be expressed as

$$P^h(99) = \widehat{W}_F^h(99)P_F^h(99) + \left(1 - \widehat{W}_F^h(99)P_{NF}^h(R,99)\right) \qquad (2)$$

Note that the weights applied to food and non-food also vary across households. The weight for each household was calculated from the predicted value of the regression of household food share in 1999, $\widehat{W}_F^h(99)$, on the logarithm of per capita consumption, and the logarithm of household size. In this manner the influence of household specific unobserved components or tastes on the share of food is eliminated.

To provide more concrete evidence about the relative price regimes prevailing in the two survey years, Figure 1 presents the change in mean prices per 1,000 calories (kcal) paid by households between 1996 and in 1999 in rural and urban areas in the province of Central Java, a densely populated province with a high concentration of poor people. The prices per 1,000 calories are calculated by dividing the nominal value of household consumption for each food group by the total quantity of kilocalories provided by all the food items in the group divided by 1000[16]. Poorer households may consume food items of lower quality and as a consequence the prices of calories paid by these households may be lower than those paid by richer households. In order to investigate for this possibility prices per 1,000 calories are also calculated separately for households at the bottom and at the top 25th percent of the distribution of total consumption per capita in each year[17].

Figure 1

Figure 1 confirms that the relative prices faced by households changed considerably between 1996 and 1999[18]. In general, the absolute price of calories from all food groups seems to have increased dramatically between the two years in both urban and rural areas. Relative to cereals, other food groups such as meat and fish, fruits and vegetables and dairy products became more expensive irrespective of whether they lived in rural or 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 and particularly so in urban areas. For example, in urban areas, the price of eggs and milk relative to cereals increased by 23 percent for the poorest consumers while the increase was 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. In Figure 2, we present the change in average daily calories per capita between 1996 and 1999. There was a significant reduction in total caloric availability and in proportional terms. Calorie availability at the household level in Central Java declined by 8 percent and 6 percent in rural and urban areas respectively. Looking across the major food groups, there was a much larger reduction in calories derived from food groups that are richer in micro-nutrients, namely, 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

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[16] 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 manage to substitute away from more expensive food items within and between groups.
[17] In 1999, the percentiles of total consumption per capita are estimated after dividing consumption by the deflator discussed earlier.
[18] For a related analysis of the impact of the Indonesian crisis on budget shares with repeated observations on sampled households see Thomas et al. (1999).

larger reductions witnessed for the poor in almost all cases, both in rural and urban Central Java. The case of calories sourced from cereals and tubers is interesting. The amount of calories obtained from cereals and tubers declined by 12 percent among the poorest in rural Central Java while the corresponding decline for the richest households was only 1 percent. In contrast, roughly the same level of declines were witnessed by richest and the poorest urban consumers. This discrepancy is, to a large extent, attributable to the fact that the rich consumers in rural areas are likely to be land owners, possibly engaged in the production of some of these staples and price increase of this nature may actually make them better off and not affect their consumption of cereals by too much.

Figure 2

The reduction in the consumption of micro-nutrient rich food groups such as meat and fish, fruits and vegetables and eggs and milk is further highlighted in Figure 3 in which we show the relative proportion of calories obtained from these food groups in 1996 and 1999. Among all groups – rich or poor – and irrespective of their residence, urban or rural, there is a marked decline in the share of calories obtained from these food groups. Compared to rural consumers, urban consumers appear to have a larger proportion of calories derived from these food groups. In 1996, the richest households in urban areas derived 6 percent of their calories from meat and fish products, 5 percent from fruits and vegetables and 4 percent from dairy products. In the same year the richest rural consumers, on the other hand, derived on average 4 percent of calories from meat and fish, 6 percent from fruits and vegetables and 2 percent from dairy products. But in the crisis year, there is also a larger reduction for urban consumers. This is, in many ways a result of the fact that urban residents generally tend to be net consumers of food and are more vulnerable during times of price increases.

Figure 3

3. Empirical Analysis and Results

Economic theory provides little guidance on how the income elasticity of any given commodity may change as a result of changes in prices. Given a Marshallian demand function for any food item x_i, summarized by the function, $x_i = x(\bar{p}, M, Z)$, where $\bar{p}$ 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 on the same set of variables, i.e.,

$$\frac{\vartheta x_i}{\partial M}(\bar{p}, M, Z).$$

The quadratic Engel curve proposed by Banks, Blundell, and Lewbel (1997) is the only study that we could identify in the literature recognizing explicitly the dependence of the income elasticity to prices. In the QUAIDS specification they propose, intended primarily at allowing the income effect to vary across different points of the income distribution, the income elasticity varies with prices but only through the coefficient of the quadratic term of income (see eq. 11 in their paper). Unless one is willing to make strong (if not arbitrary) assumptions about the separability of preferences between and within specific food groups there is little that can be said (at least from a theoretical perspective) about how changes in the prevailing relative prices may affect the response of demand for a commodity to income changes. Under these circumstances, this issue can only be addressed empirically. This is precisely the gap that our study aims to fill. The question we address empirically is whether the income elasticity differs significantly between a non-crisis year (1996) and a crisis year (1999) characterized by very different relative price vectors, $\bar{p}_{96}$ and $\bar{p}_{99}$, respectively. Based on the notation above, we investigate whether

$$\frac{\vartheta x_i}{\partial M}(\bar{p}_{96}, M, Z) \neq \frac{\vartheta x_i}{\partial M}(\bar{p}_{99}, M, Z)$$

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

3.1 Nonparametric Regression – The Starchy Staple Ratio

We begin our analysis with an investigation of whether and how the Starchy Staple Ratio (SSR) and its sensitivity to income varies in a cross section of households in 1996 to 1999. The SSR, calculated as the share of total calories obtained from cereals and tubers, is considered to be a more useful aggregate measure of household welfare than total caloric availability per capita since it captures the diversity in dietary patterns. According to Bennett's Law, the SSR or the share of dietary calories obtained from cereals and tubers declines with the level of household income. [19] We adopt a flexible approach that simply examines for Indonesian households by aggregating the calorie contents of the more than 200 food items included in the SUSENAS survey. Given that the calorie income elasticity in Indonesia is known to be nonlinear (Skoufias, 2003), we use nonparametric methods that also allow the income elasticity of the SSR to vary with the level of income. Using y to denote the logarithm of the SSR, and x the logarithm of per capita total household consumption, the regression function, can be written as

$$m(x) = E(y|x) \tag{3}$$

Following Subramanian and Deaton (1996) and Deaton (1997), we estimate $m(x)$ using a smooth local regression technique proposed by Fan (1992)[20]. At any given point x, we run a weighted linear regression of the logarithm of the SSR on the logarithm of per capita consumption. The weights are chosen to be largest for sample points close to x and to diminish with distance from x. Instead of estimating a regression for each point x in the sample, we divided the distribution of log per capita consumption into 100 evenly spaced grids and estimated local regressions for each grid. For the local regression at x, observation i gets the (quartic kernel) weight

[19] Bennett's law is about the relation between household diet and income whereas Engel's law relates to the share of of food expenditures in the household budget and the level of household income. Timmer, Falcon and Pearson (1983) provide a more detailed discussion of Bennett's law.

[20] Fan (1993) has demonstrated the superiority of the smooth local regression technique over kernel and other methods.

$$w_i(x) = \frac{15}{16}\left[1 - \left(\frac{x - x_i}{h}\right)^2\right]^2 \quad (4)$$

if $-h \leq x - x_i \leq h$ and zero otherwise. The quantity h is a bandwidth that is set so as to trade off bias and variance, and that tends to zero with the sample size. We have set the bandwidth to the value of 0.8.

An advantage offered by these nonparametric graphs is that potential biases in the measurement of the level of calorie availability among households of higher level of income do not impact on estimates of the elasticity among poorer households. Inspection of the nonparametric graphs, for example, allows one to obtain a sense of the extent to which the exclusion of nutrients obtained from prepared foods, which is more common among wealthier households, affects the nutrient income elasticity of wealthier households.

The left panels of Figure 4a and 4b below present graphs of the log of SSR of the dietary content in 1996 and in 1999 against the log of per capita consumption (in 1996 prices) for rural and urban areas in the province of Central Java. Given that the level of income in 1999 has been made comparable to that in 1996, these graphs provide an opportunity to examine the effect of relative price changes between 1996 and 1999 on household dietary composition as summarized by the SSR assuming that family size has not changed significantly over the same period.

For rural areas in Central Java, the SSR in 1999 appears to pivot slightly just below the median level of per capita expenditures in rural Central Java (denoted by the vertical line in the graph) with the level of the SSR being slightly below that of 1996 for the relatively poorer households and higher than the 1996 level of SSR for the households in the upper half of the distribution. This suggests that in the crisis year the availability of calories derived from cheaper food sources such as cereals and tubers seems to have generally increased across the relatively higher income groups in rural areas in Central Java.[21] The pattern is even more obvious in the urban areas where the SSR line in 1999 lies almost uniformly above the 1999 SSR line. Thus, for the urban poor, there is a rather clear shift to the right in 1999 indicating a larger reliance on the starchy staples for calories in times of crisis. Among rural households the pattern is not so clear.

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[21] The increase in household food insecurity and compromised diet in three Java provinces, including Central Java, during the crisis is also corroborated by Studdert et al. (2001).

The differences in the slope of the SSR lines in 1999 compared to the respective 1996 lines in both rural and urban areas, suggest that the responsiveness of the SSR to increases in income varies across the non-crisis and crisis years. The right panels of Figures 4a and 4b present graphs of the income elasticity of the SSR in 1996 and in 1999 against the log of per capita consumption (in 1996 prices) for rural and urban areas, respectively. In the rural areas, the elasticity of the SSR with respect to income is higher during the crisis year and invariant to the level of household income at about -0.27 percent. Thus, it appears that an increase in income, such as that resulting by a program of cash transfers to poorer rural households during a crisis year is likely to be more effective in increasing dietary diversity (i.e. decreasing the SSR) compared to a non-crisis year. In contrast, in the urban areas where there is a larger reliance on starchy staples for calories in the crisis year, the income elasticity of the SSR appears to be about the same in 1996 and 1999 for the lower levels of income, becoming smaller than the level of the elasticity in the non-crisis year.[22]

Figures 4a & 4b

3.2 Regression Analysis

The analysis in the previous section focused on the bivariate relationship between SSR and total per capita expenditures. While being informative about the general shape of the relationship and how it changed in the two years, these graphs are unable to account for a number of critical factors, the primary among these being the differences in the age and gender composition of households, and the problem of correlated errors between household consumption and nutrients. In this section, we examine specifically the elasticity of the SSR with respect to household consumption controlling for the usual set of household characteristics.

[22] InAnnex A we examine whether the elasticity estimates are significantly different at different levels of outlay by checking whether the standard error bands for the 1996 estimate overlaps with the standard error bands for the 1999 estimate. These graphs do not reveal any significant differences in the income elasticity estimates during the crisis and no-crisis years, a result that is also confirmed with the regression methods employed in the latter part of the paper.

The focus on the impacts of the relative prices vectors on the income elasticity of demand faces some constraints.[23] In any given year, a typical cross-sectional household survey collects data within a short time interval. As a consequence, most of the variation in the price of any given commodity faced by households arises from differences in the quality of the commodity consumed, transportation costs, market segmentation and other transaction costs that may prohibit the equalization of consumer prices across space. To the extent that households in different locations are surveyed in different quarters in the calendar year then the survey may also capture seasonal variability in prices. But even if this were possible, it is still doubtful whether the seasonal variation in prices provides an adequate representation of the change in relative prices that consumers face during a major economic crisis. Household panel data provide an opportunity to relax some of these shortcomings. Behrman and Deolalikar (1987), for example, analyze the calorie income relationship using data from the Village Level Survey of ICRISAT. But even these data can shed little light on this question since there was a relatively stable economic environment during the period of the study.

In order to compensate for the fact that our data are not longitudinal, but rather cross-sectional surveys applied in different years, we adopt a flexible specification that provides an explicit test of the difference between the elasticity coefficients in the pre-crisis and crisis years. In order to implement this, we pool the two cross-sectional data and run the following regression:

$$\begin{aligned} \ln(\mathrm{SSR}_{\mathrm{jkt}}) = {} & (\alpha_{96} + \beta_{96}\mathrm{lnPCE}_{\mathrm{jkt}} + \gamma'_{96}\mathrm{X}_{\mathrm{jkt}} + \mu_{96}) + \\ & + \mathrm{D}_{99} * (\alpha_{99} + \beta_{99}\mathrm{lnPCE}_{\mathrm{jkt}} + \gamma'_{99}\mathrm{X}_{\mathrm{jkt}}) + \mu_{99} + \varepsilon_{\mathrm{jkt}} \end{aligned} \tag{5}$$

where $lnSSR_{jkt}$ is the log of SSR for household j that lives in cluster k in year t, μ_{96} and μ_{99} are the vectors of binary variables identifying the cluster fixed effects in the 1996 and 1999 rounds[24], and D_{99} is binary dummy variables equal to 1 for observations in 1999 and equal to 0 otherwise.

[23] In much of the literature on the nutrient-income relationship (e.g. see Strauss and Thomas, 1995) prices are typically left out of the specification of the Engel curve estimated with 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 et al. (1997).

[24] The SUSENAS survey is a clustered survey with at most 16 households surveyed per cluster in each year. Clusters in 1996 and in 1999 have the same code, but given that we were unable to ascertain whether they represented the same clusters across the two survey years we treated clusters with the same code in different years as different clusters.

The variable $lnPCE_{jkt}$ denotes real per capita consumption expenditures. The set of control variables X is a vector of household characteristics and ε_{jkt} is an error term summarizing the influence of random disturbances. The elements of the vector X are as follows: 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 0-5 years of age, the number of children 6-12 years of age, the number of males and females 13-19 and 20-54 years of age and the number of males greater than 55 years of age). The list of additional binary variables includes whether the household head is a female, and a group of a group of dummy variables describing the educational level of the household head and his/her spouse, such as whether he/she completed primary school, junior high school, or senior high school, and the sector of employment of the household head and his/her spouse, such as whether one is self-employed, unemployed or wage worker.

With this specification, we control for the different relative prices prevailing across the two survey years through the cluster-level fixed effects that are allowed to differ across the two survey years.[25] In addition, we allow the coefficients of all the control variables as well as the coefficient of $lnPCE_{jkt}$ to differ across the two years. This provides us with an explicit test of the difference in the income elasticity of the various dependent variables between 1996 and 1999. The dummy variable D_{99} included in the regression is also able to absorb any other aggregate effect that may have changed in addition to relative prices between 1996 and 1999. Even though the cluster fixed effects in this specification absorb the prices of food items, which vary only at the cluster level, we also estimate a similar equation controlling specifically for the prices of the five aggregate food groups.[26].

As first pointed out by Bouis and Haddad (1992), a food expenditure survey may overstate the nutrient availability within wealthier households since it is common for these households to provide meals to employees and domestic servants[27]. In addition, following the 1997-98 crisis it is plausible that there is an increase in the frequency of this practice. To

[25] The same approach is also employed with cross-sectional data by Subramanian and Deaton (1996). Cluster fixed-effects also take into account other time-invariant local characteristics that determine dietary intakes and preferences.
[26] In order to calculate the price of each of the food groups, we first construct unit values of each of the food items that constitute the group for every household. Unit values are of course a biased measure of prices as they are contaminated with unobserved quality of the food items consumed. We average the unit values across households within each village and consider the average the price faced by each household within the village.
[27] It should be noted that in the SUSENAS, domestic servants are counted as members of the household roster.

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 is done by limiting the sample to the lower half of the distribution of consumption per capita within rural and urban areas separately within the province of Central Java. Robust standard errors are estimated in order to control for unknown forms of heteroskedasticity.

In defense of the OLS estimates, it is important to bear in mind that the SUSENAS survey is a 7-day food intake or consumption survey that carefully collects information about food consumed outside the household or food received in kind from sources outside the household. Receipts of specific food items as payment for services provided or food gifts and transfers are captured by the survey. In addition, since the exact same questionnaire was applied at the same point in time in each survey year, there is no reason to believe that biases may arise due to these factors.

Yet, the possibility remains that correlated measurement errors in the total food consumption and thus caloric and nutrient availability are one potential source of bias in estimates of nutrient income elasticities. As first noted by Bouis and Haddad (1992) in the linear version of equation (3) the possibility that measurement errors in nutrient availability are likely to be 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 towards 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 will outweigh 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 (3), Deaton notes that elasticity estimates using the log of non-food expenditures as the sole instrumental variable are likely to be biased downward, implying that the combination of the elasticities estimated by OLS and instrumental variable methods provide upper and lower bounds, respectively, for their true value. In order to address these considerations we used an index of household assets constructed using the principal components method, as an instrumental variable for $lnPCE_{jkt}$.

Specifically, in each survey year we estimated the index of household assets using the following variables: number of cows and buffaloes, sheep and goats, pigs, chicken and ducks owned by the household as well as dummy variables summarizing the household residence and its environment, such as roof made of concrete or tile, walls made out of brick or wood, floor made out of tile or cement, a private or shared toilet, drinking water access to public network or pump or purchased and whether energy for cooking, lighting and heating is obtained through the public electric network or gas. Given that the consumption variable interacted with the 1999 year dummy enters the estimated equation, the instrument we used for this interaction term is the asset index based on 1999 data interacted with the 1999 year dummy. We present the results of the first stage regressions in Annex B. [28]

In Table 1, we present the elasticity estimates for SSR obtained from these regressions. Model I corresponds to the specification in which we don't control for the prices while in Model II, we specifically control for the prices of the main food groups. The estimates for these two models reveal that we get the same results irrespective of whether we control for prices explicitly or let the fixed effects absorb the influence of prices. As we would expect, the negative sign on the point estimates for the income elasticity of SSR is consistent with the non-parametric 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 the SSR elasticity is marginally higher in rural Central Java. This implies that the 'rate' at which households switch from cheap sources of calories to food groups that are better sources of micro-nutrients as their incomes rise is higher in rural areas than in urban areas.

The estimated coefficient for β_{99} in our regression is of particular interest because it contains information on how different the elasticities were during the 1999 crisis year, in comparison to the reference year of 1996. The estimates for β_{99} in the IV specification reported in Table 1 suggest that the SSR elasticity in 1999 was practically identical to the elasticity in 1996. The size of the marginal effect is very small and it is not statistically significant. This implies that the income elasticity of SSR, which is a measure of the degree to which households

[28] In all cases, a Hausman-type test (Holly, 1982) for the absence of measurement error in the consumption variable rejected the null hypothesis. We also rejected a similar test for the exogeneity of the per capita consumption and its interaction in equation 3.

diversify away from starchy staples was invariant to the different relative prices prevailing during the crisis year.

TABLE 1

Even though the invariance of the SSR elasticity to price increases is suggestive of little substitution towards or away from the cereals and tubers category, the complex pattern of price changes witnessed in Indonesia during the financial crisis could well have induced households to make substitutions within and among other food groups. In order to probe deeper into the consequences of such substitutions for the elasticity of micro-nutrients, we repeat the estimation embodied in the equation (3) but with the log of the consumption of specific nutrients as dependent variables. The elasticity coefficients obtained for these regressions are reported in Table 2.

The OLS estimates of the nutrient elasticities for both rural and urban Central Java are all positive and statistically significant in the reference year. For example for rural areas, the estimates for the elasticity range from 0.18 for carbohydrates to 0.63 for fat in the reference year of 1996. For the urban areas, the spread is narrower and ranges from 0.13 for carbohydrates and 0.44 for fat. The IV estimates, on the other hand, while generally statistically significant, appear to be smaller in magnitude in comparison to the corresponding OLS estimates both urban and rural Central Java. 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 the only nutrient for which the IV estimates of the elasticity are not significant in both urban and rural areas is Vitamin C. This is suggestive of the possibility that Vitamin C consumption may not be responsive to income in Central Java, irrespective of whether it's a crisis year or not.

Second, the estimates for β_{99}, which is the coefficient on the interaction between the log of per capita expenditure and the year dummy for 1999, appear statistically significant for calories, protein, fat, carbohydrates, phosphorous, iron, and Vitamin B for rural areas and calories, protein, fat, calcium, phosphorous and iron for urban areas. This is evidence of the elasticity of these nutrients being different in the crisis year in comparison to the reference year. The magnitudes of these differences are quite large as well, particularly for urban areas. The income elasticity of calories and iron appears to have doubled in urban areas in 1999. The elasticities for protein, calcium and phosphorous also came close to doubling. The income elasticity of fat appears to have more than doubled in 1999 for urban households.

However, there are some nutrients for which the income elasticity did not change significantly between 1996 and 1999. These are calcium and vitamin A in rural Central Java and carbohydrates, vitamin A and vitamin B in urban Central Java. An increase in income elasticity of any particular nutrient in the crisis year – particularly of the magnitudes witnessed in urban areas – may be considered as indicating a deterioration of the nutritional status of households. Consider the income elasticity for fat which is generally the highest for both rural and urban consumers. Among poor households, it reflects the relative 'luxury good' status of the food items such as meat and fish, from which it is primarily sourced. The price increases in 1999 appear to have severely curtailed households' ability to access basic nutrients such as calories, protein, calcium, phosphorous and iron.

TABLE 2

4. Conclusion

In this paper we have examined the size of the income elasticity of the demand for nutrients and its sensitivity to changes in the relative prices and the economic environment faced by households. Using household consumption data from the 1996 and 1999 consumption module of the National Socio-Economic Surveys (SUSENAS) in Indonesia our analysis revealed that there is considerable heterogeneity in the income elasticity of demand for nutrients over time and depending on the econometric method used. A comparison of the OLS and instrumental variable estimates of the demand for nutrients suggests that the OLS estimates are likely to be misleading due to the bias from the correlated errors in consumption and nutrient content. In particular, the finding that the instrumental variable estimates are generally lower in magnitude than their OLS counterparts, suggest that the upward bias due to correlated measurement errors in the nutrient intakes and household level consumption may outweigh the possible attenuation bias.

We also find that for most nutrients, including micro-nutrients such as phosphorous, iron and calcium, the income elasticity estimates are significantly higher in the crisis year.

Moreover, the magnitude of the increase appears to be generally larger in urban areas than in rural areas. On the other hand, for some nutrients such as vitamin C, the income elasticity estimates obtained from the instrumental variables specification appear to be statistically indistinguishable from zero. This suggests that income may have limited leverage to increase or protect the consumption of vitamin C irrespective of whether it is a crisis year or not.

Our results have two specific implications for policy. First, given the significant increases in the income elasticity of both micro and macro-nutrients during the times of a crisis, it appears that cash transfer programs can play an integral role in helping households protect their consumption of essential nutrients during crisis. To the extent that the delivery infrastructure already exists and operates at low levels of leakage, cash transfers are widely accepted as the quickest and the cheapest interventions to scale up in order to reach households that are most likely to be adversely affected. This research shows that they can also be more effective in protecting the consumption of nutrients than in normal economic conditions.

The second implication that emerges from this paper is that if the goal of policy in response to economic crises is to ensure the protection of all important micro-nutrients, then relying entirely on cash transfers may not be sufficient. For example, the consumption of an important micro-nutrient, vitamin C, was found not to be responsive to income in both rural and urban Central Java. This suggests that specifically targeted micro-nutrient supplementation programs may have to accompany cash transfers to ensure that key micro-nutrients are not sacrificed during crises. Future research could be directed at understanding and identifying which specific nutrients households are likely to sacrifice during a crisis in different settings.

REFERENCES

Banks, J. R. Blundell, and A. Lewbel (1997) "Quadratic Engel Curves and Consumer Demand," *The Review of Economics and Statistics,* Vol. LXXIX no. 4, (November), pp. 527-39.

Behrman, J., (1995). "Micronutrients and Household Behavior: What we Know and What we Don't Know." Working Papers Series on Agricultural Strategies for Micronutrients, Number 2, International Food Policy Research Institute, Washington DC.

Behrman, J. (1988). "Nutrition and Incomes: Tightly Wedded or loosely Meshed?" Pew/Cornell Lecture Series on Food and Nutrition Policy. Ithaca: Cornell University.

Behrman, J., and A. Deolalikar. (1989). "Is Variety the Spice of Life? Implications for Calorie Intake." *The Review of Economics and Statistics*, vol. 71, no. 4, pp. 666-672.

Behrman, J., and A. Deolalikar. (1987). "Will Developing Country Nutrition Improve With Income? A Case Study for Rural South India." *Journal of Political Economy,* vol. 95, pp. 492-507.

Behrman, J., A. Foster, and M. R. Rosenzweig (1997). "The Dynamics of Agricultural Production and the Calorie-Income Relationship: Evidence from Pakistan," *Journal of Econometrics*, vol. 77, pp. 187-207.

Behrman, J. and B.L. Wolfe (1984) "More Evidence on Nutrition Demand: Income Seems Overrated and Women's Schooling Underemphasized," *Journal of Development Economics*, vol. 14 (1-2), pp. 105-126.

Bouis, H., (1991). "The Determinants of Household-level Demand for Micronutrients: An Analysis for Philippine Farm Households." International Food Policy Research Institute. Washington, D.C. Mimeo.

Bouis, H., and L. Haddad. (1992). "Are Estimates of Calorie-Income Elasticities too high? A Recalibration of the Plausible Range." *Journal of Development Economics*, vol. 39, no. 2, pp.333-364.

Chernichovsky, D., and O. Meesook. (1984). "Urban-rural food and nutrition consumption patterns in Indonesia." World Bank Staff Working Paper No. 670. Washington D.C.: World Bank.

Cramer J. C. (1973) "Interaction of Income and Price in Consumer Demand," *International Economic Review*, Vol 14, (June), pp. 351-363

Deaton, A. (1995). "Data and Econometric Tools for Development Analysis," chapter 33 in *Handbook of Development Economics* vol. 3, eds. J. Behrman and T.N. Srinivasan. Amsterdam: North-Holland.

Deaton, A. 1997. *The Analysis of Household Survey Data*. Baltimore: Johns Hopkins University Press.

Deaton, A., and S. Zaidi. (1999). Guidelines for Constructing Consumption Aggregates for Welfare Analysis. Princeton University. Mimeo

Direktorat Gizi, Departemen Kesehatan (1988) *Daftar Komposisi Bahan Makanan*. Jakarta: Penerbit Bhratava karya Aksra.

Efron, B., and Tibshirani, R.J. (1993). *An introduction to the bootstrap*. London: Chapman and Hall.

Fan, J. (1993). "Local Linear Regression Smoothers and their Minimax efficiencies." *Annals of Statistics*, vol. 21 (March), pp. 998-1004.

Hausman, J. A. (1978). Specification tests in econometrics. *Econometrica*, 46: 1251-1271.

Holly, A. (1982). A Remark on Hausman's Specification Test," *Econometrica*, vol. 50, pp. 749-759.

Ivanic, M and W. Martin (2008) "Implications of higher global food prices for poverty in low-income countries" *Agricultural Economics*, International Association of Agricultural Economists, vol. 39(s1), pages 405-416, November.

Jensen, Robert T., and Nolan H. Miller. 2008. "Giffen Behavior and Subsistence Consumption." *American Economic Review*, 98(4): 1553–77.

Levinsohn, J., S. Berry, and J. Friedman (1999), "Impacts of the Indonesian Economic Crisis: Price changes and the Poor," NBER Working Paper No. 7194, June, National Bureau of Economic Research, Cambridge, MA.

Molyneaux J. (2002) "Starchy Staple Consumption and Household Welfare: A Fresh look at Indonesian Welfare," DAI-Food Policy Support Activities/MOA-Food Security Agency.

Pitt, M. M. (1983) "Food Preferences and Nutrition in Rural Bangladesh," *The Review of Economics and Statistics*, vol. 65, no. 1, pp. 105-114.

Pitt, M. M. and M. R. Rosenzweig (1985) "Health and Nutrient Consumption Across and Within Farm Households." *The Review of Economics and Statistics*, vol. 67 (2), pp. 212-223.

Ravallion, M. (1990). "Income Effects on Undernutrition." *Economic Development and Cultural Change,* vol. 38 pp. 489-515.

Skoufias, Emmanuel (2003) "Is the Calorie-Income Elasticity Sensitive to Price Changes? Evidence from Indonesia," *World Development*, (July) Vol. 31, no. 7.

Skoufias, Emmanuel (2001) "Changes in Regional Inequality and Social Welfare in Indonesia from 1996 to 1999." *Journal of International Development*, Vol. 13, No. 1, (January), pp. 73-91.

Strauss, J., and D. Thomas. (1995). "Human Resources: Empirical Modeling of Household and Family Decisions." Chapter 34 in *Handbook of Development Economics* vol. 3, eds. J. Behrman and T.N. Srinivasan. Amsterdam: North-Holland.

Studdert, L. E. Frongillo, and P. Valois (2001) "Household Food Insecurity was Prevalent in Java during Indonesia's Economic Crisis," *Journal of Nutrition*, vol. 131, no. 10, pp. 2685-2691

Subramanian, S. and A. Deaton. (1996). "The Demand for Food and Calories." *Journal of Political Economy*, vol. 104, pp. 133-162.

Suryahadi, A., S. Sumarto., Y. Suharso, and L. Pritchett (2000), "The Evolution of Poverty During the Crisis in Indonesia, 1996 to 1999 (Using Full SUSENAS Sample)." SMERU Working Paper, March, Social Monitoring & Early Response Unit, Jakarta.

Thomas, D., E. Frankenberg, K. Beegle, and G. Teruel (1999), "Household Budgets, Household Composition and the Crisis in Indonesia: Evidence from Longitudinal Household

Survey Data," Paper prepared for presentation at the 1999 Population Association of America Meetings, RAND, Santa Monica, CA.

Timmer P. (1981) "Is There Curvature in the Slutsky Matrix?" *The Review of Economics and Statistics*, Vol. 62, no. 3, (August), pp. 395-402

Timmer, P., and H. Alderman. (1979). "Estimating Consumption Parameters for Food Policy Analysis." *American Journal of Agricultural Economics*. vol. 61, pp. 982-994.

Timmer P., W.P. Falcon, S. Pearson (1983) *Food Policy Analysis* Published for the World Bank. Johns Hopkins University Press

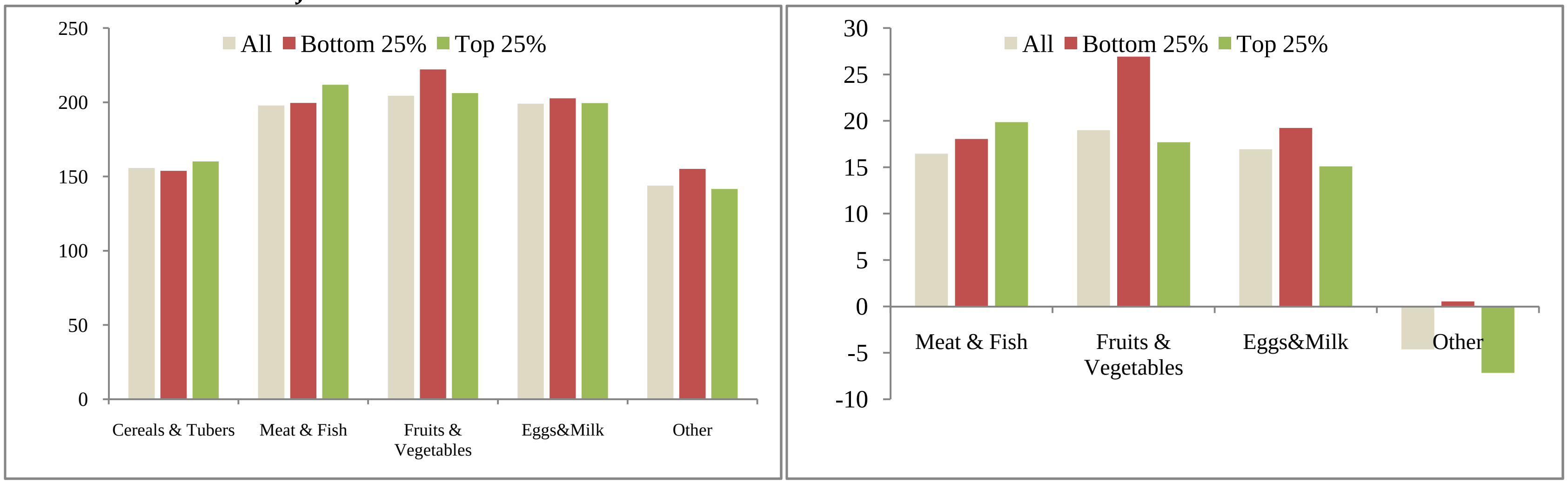
Tiwari S., and H. Zaman (2010), "The Impact of Economic Shocks on Global Undernourishment" Policy Research Working Paper 5215, The World Bank

Van de Walle, Dominique (1988), "On the Use of the SUSENAS for Modeling Consumer Behavior," *Bulletin of Indonesian Economic Studies*, Vol. 24, No. 2, (August):, pp. 107-122.

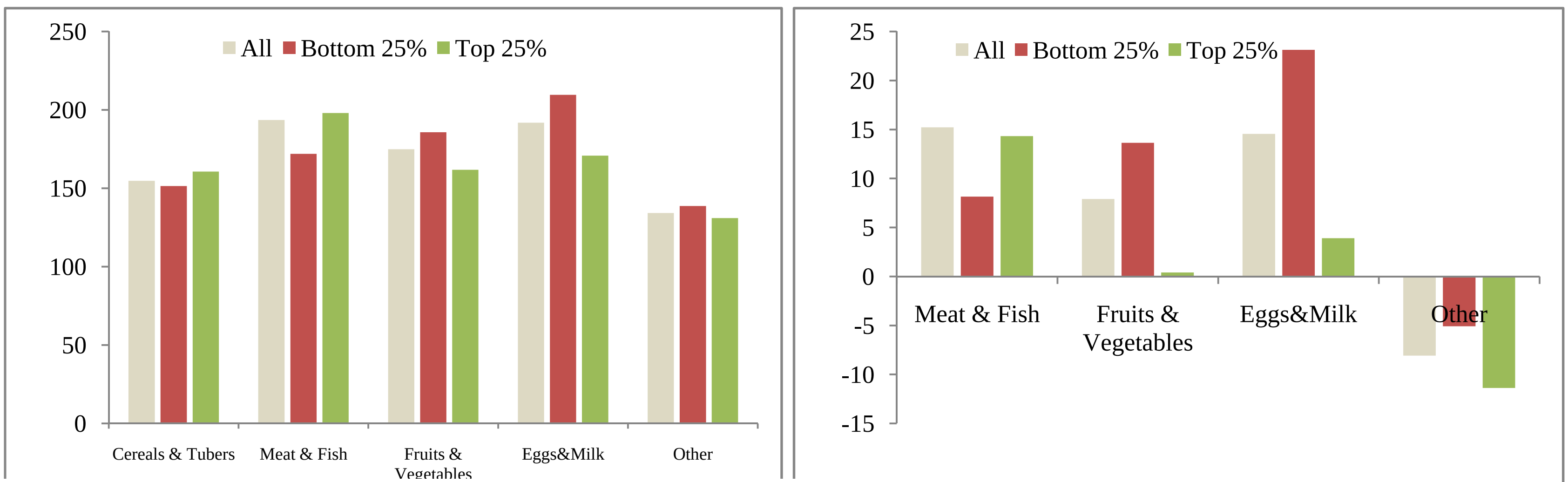
Wodon, Q., and H. Zaman (2010), "Higher Food Prices in Sub-Saharan Africa: Poverty Impact and Policy Responses", *World Bank Research Observer Vol. 25 (1)*

Figure 1: Change in the absolute and relative price of 1000 calories from 1996 to 1999, by food groups.

Panel A: Rural Central Java



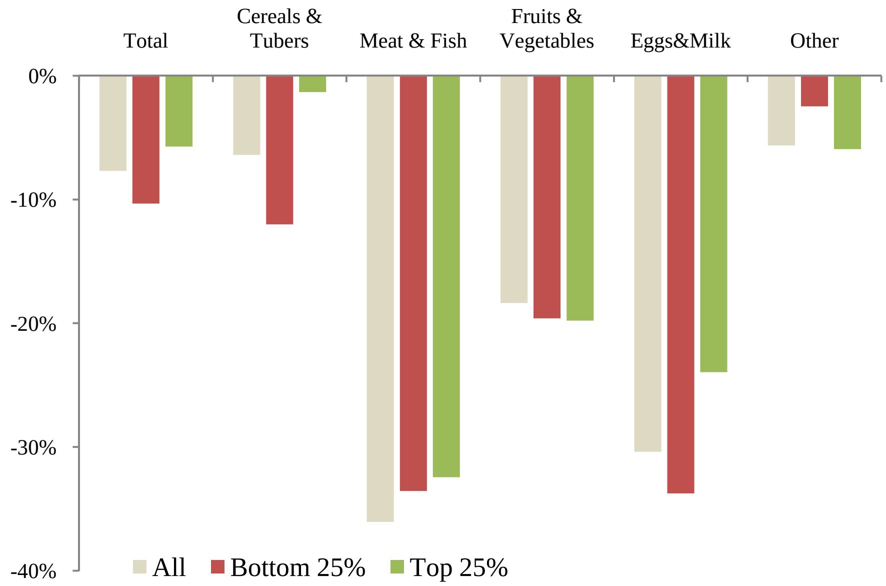
Panel B: Urban Central Java



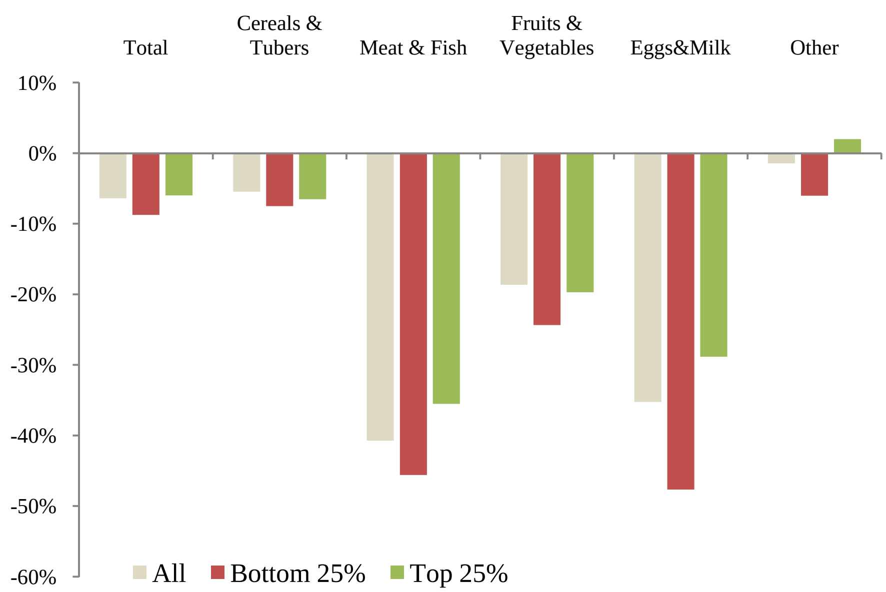
Source: Authors' calculations based on 1996 and 1999 SUSENAS consumption modules. The charts on the left panel show the change in the price of 1000 calories in 1999 from the respective calorie price in 1996 and charts on the right show the change in the relative calorie price between 1996 and 1999 (relative to the calorie price from cereals & tubers in 1996 and in 1999, respectively). Prices are reported separately for the top and bottom quintile of the expenditure distributions.

Figure 2: Change in per capita calorie consumption, by food groups

Panel A: Rural Central Java



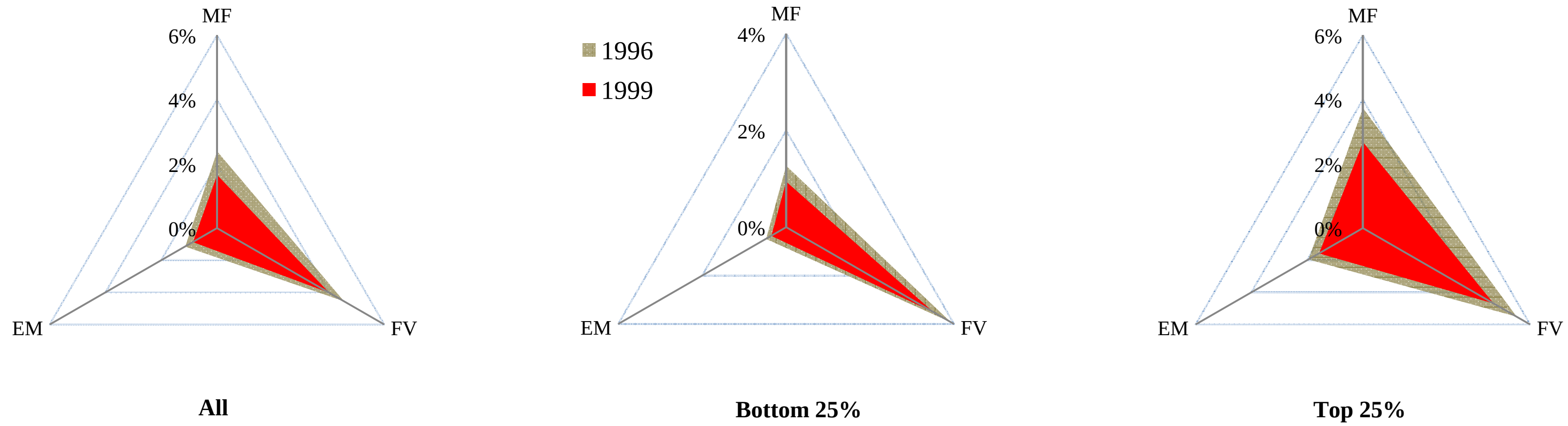
Panel B: Urban Central Java



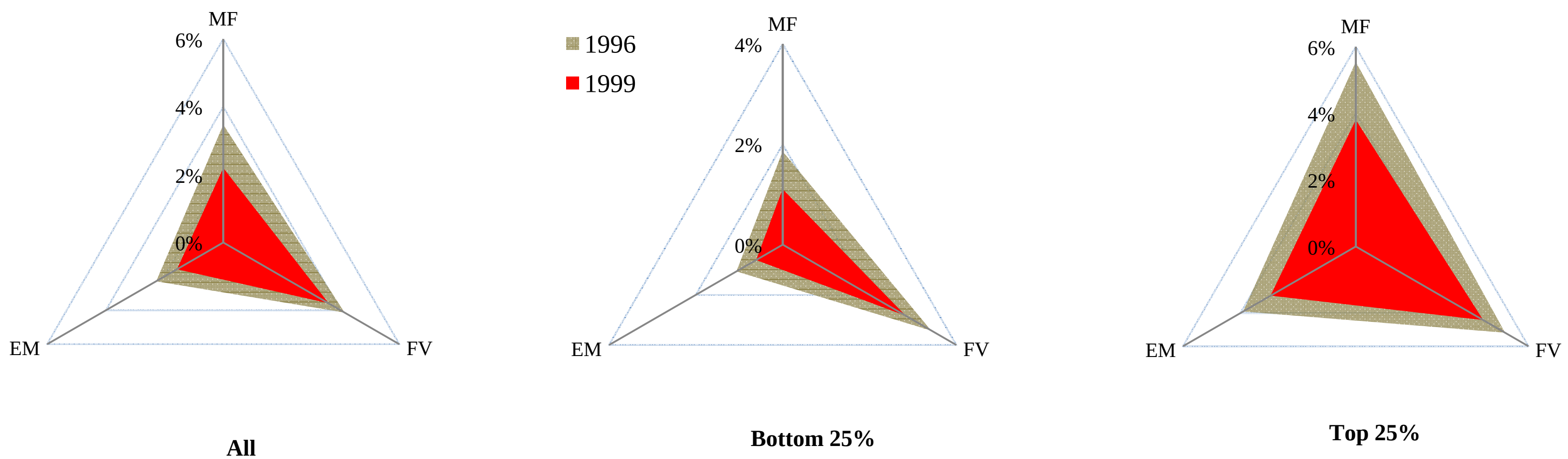
Source: Authors' calculations based on 1996 and 1999 SUSENAS consumption modules. Changes in calorie consumption between 1996 and 1999 are reported separately for the top and bottom quintile of the expenditure distributions.

Figure 3: Calorie share of micronutrient rich food groups in 1996 and 1999

Panel A: Rural Central Java (All, Bottom 25% and Top 25%)



Panel B: Urban Central Java (All, Bottom 25% and Top 25%)

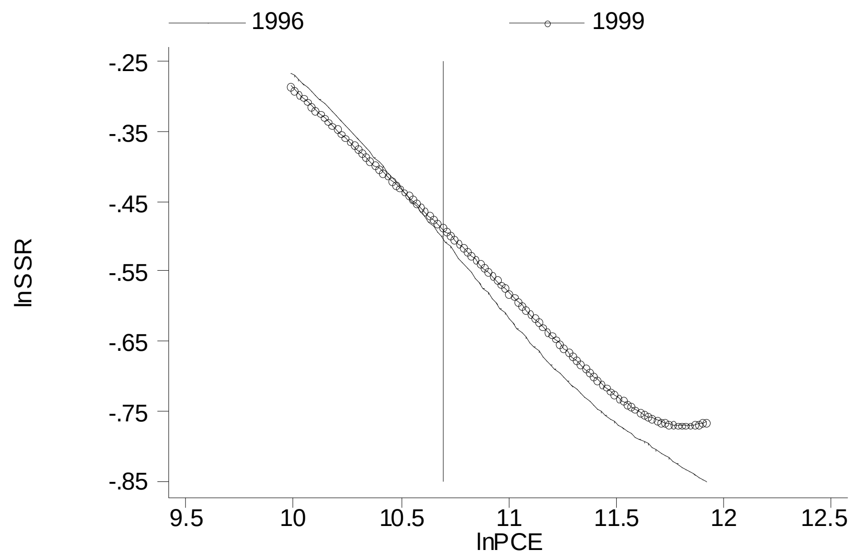


NOTE: Authors' calculations using SUSENAS consumption data for 1996 and 1999. MF denotes Meat and Fish, FV denotes fruits and vegetables and EM denotes Eggs and Milk. For each panel, the first chart corresponds to all households, the second chart for the bottom quintile of the expenditure distribution and the third chart to the top quintile.

Figure 4: Some characteristics of the starchy staple ratio in Urban and Rural Central Java, 1996 and 1999

A: Rural Central Java

The relationship between lnSSR and lnPCE



The income elasticity of SSR against lnPCE

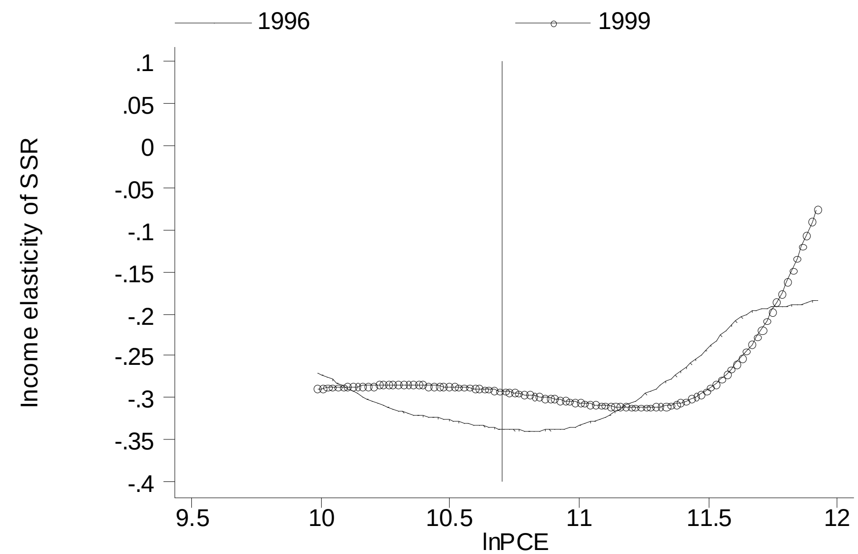
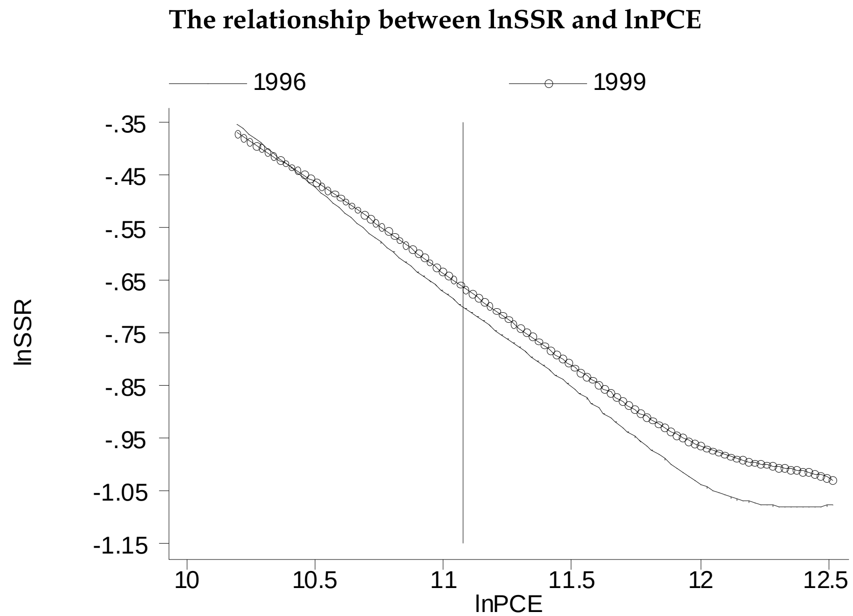


Figure 4: Some characteristics of the starchy staple ratio in Urban and Rural Central Java, 1996 and 1999

B: Urban Central Java



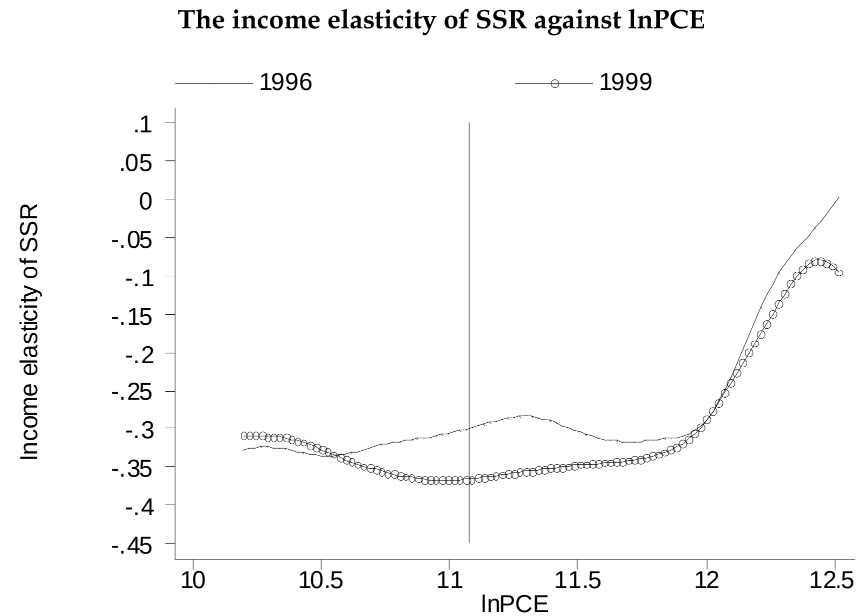


Table 1: Calorie share elasticity estimates for Poor Households in Central Java

	Model I		Model II	
VARIABLES	OLS	IV	OLS	IV
PANEL A: Rural Central Java				
log PCE	-0.25***	-0.24***	-0.25***	-0.24***
	(0.01)	(0.01)	(0.01)	(0.01)
Marginal effect (1999)[logPCE * D99]	0.01	0.00	0.01	0.00
	(0.02)	(0.02)	(0.02)	(0.02)
PANEL A: Urban Central Java				
log PCE	-0.26***	-0.22***	-0.26***	-0.22***
	(0.02)	(0.02)	(0.02)	(0.02)
Marginal effect (1999)[logPCE * D99]	-0.02	-0.00	-0.02	-0.00
	(0.02)	(0.04)	(0.02)	(0.04)

NOTES: Sample limited to the lower half of the distribution of consumption per capita within rural and urban areas separately within the province of Central Java.
Robust standard errors in parenthesis corresponding to the elasticity estimates.
Within each panel, every column represents a separate regression and a wide array of household level economic and demographic variables are controlled for.
In addition, all specifications also control for cluster fixed effects and year dummies.
Model 2 is different from Model 1 in that we control specifically for price indices of five food groups at the *kecamantan* level.
The reported IV estimates are obtained by instrumenting the log of per capita expenditure with the household specific asset index.
*** $p<0.01$, ** $p<0.05$, * $p<0.1$

Table 2: Nutrient-income elasticity estimates for Poor Households in Central Java

	Calories	Protein	Fat	Carboh.	Calcium	Phos.	Iron	Vit. A	Vit. B	Vit. C
Panel A: Rural Central Java										
OLS										
log PCE	0.25***	0.31***	0.63***	0.18***	0.37***	0.22***	0.28***	0.27***	0.19***	0.20***
	(0.02)	(0.02)	(0.04)	(0.01)	(0.03)	(0.02)	(0.03)	(0.06)	(0.02)	(0.06)
Marginal effect (1999)[logPCE * D99]	0.05**	0.09***	0.03	0.07***	0.10**	0.10***	0.16***	-0.08	0.11***	-0.07
	(0.03)	(0.03)	(0.06)	(0.02)	(0.04)	(0.03)	(0.04)	(0.08)	(0.04)	(0.09)
IV										
log PCE	0.14***	0.19***	0.50***	0.07***	0.22***	0.10***	0.16***	0.13***	0.07***	-0.03
	(0.02)	(0.02)	(0.04)	(0.02)	(0.03)	(0.02)	(0.03)	(0.05)	(0.02)	(0.05)
Marginal effect (1999)[logPCE * D99]	0.07**	0.10**	0.19***	0.06**	0.07	0.10***	0.14***	-0.15	0.13**	-0.10
	(0.03)	(0.04)	(0.07)	(0.03)	(0.05)	(0.04)	(0.04)	(0.09)	(0.05)	(0.09)
Panel A: Urban Central Java										
OLS										
log PCE	0.19***	0.26***	0.44***	0.13***	0.36***	0.22***	0.25***	0.30***	0.19***	0.31***
	(0.01)	(0.02)	(0.04)	(0.01)	(0.03)	(0.02)	(0.03)	(0.05)	(0.02)	(0.05)
Marginal effect (1999)[logPCE * D99]	0.06*	0.06	0.23***	0.04	0.11**	0.05	0.11**	0.06	0.07	0.05
	(0.03)	(0.04)	(0.06)	(0.03)	(0.05)	(0.04)	(0.05)	(0.08)	(0.05)	(0.08)
IV										
log PCE	0.12***	0.18***	0.34***	0.07***	0.26***	0.15***	0.17***	0.22***	0.13***	0.09
	(0.02)	(0.02)	(0.04)	(0.02)	(0.03)	(0.02)	(0.03)	(0.06)	(0.03)	(0.06)
Marginal effect (1999)[logPCE * D99]	0.12**	0.16**	0.38***	0.08	0.23***	0.14**	0.17**	0.07	0.08	0.10
	(0.05)	(0.07)	(0.10)	(0.05)	(0.08)	(0.07)	(0.07)	(0.12)	(0.08)	(0.11)

NOTES: Sample limited to the lower half of the distribution of consumption per capita within rural and urban areas separately within the province of Central Java. Robust standard errors in parenthesis corresponding to the elasticity estimates. Within each panel, every column represents a separate regression and a wide array of household level economic and demographic variables are controlled for. In addition, all specifications also control for cluster fixed effects and year dummies.
The reported IV estimates are obtained by instrumenting the log of per capita expenditure with household specific asset index.
*** p<0.01, ** p<0.05, * p<0.1

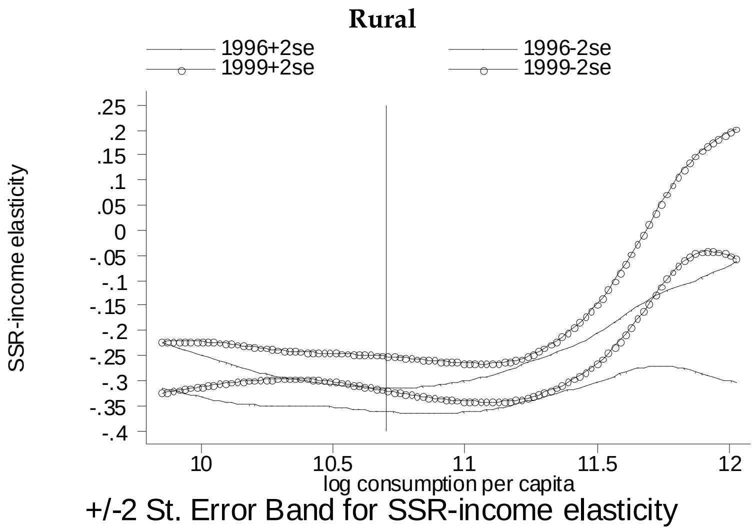
Annex A: Standard error bands for income elasticity estimates of the SSR

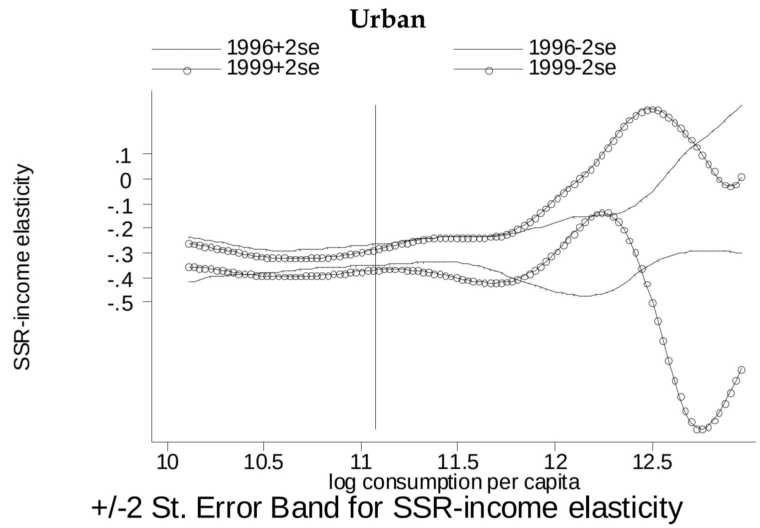
In order to determine whether the two elasticity values at each level of income are significantly different from each other it is essential to have some estimates of the standard error associated with each of the elasticity values. Figure A below contains graphs of the standard error bands separately for the 1996 and 1999 estimates of the income elasticity of the SSR. These were calculated using the formula $m'(x) \pm 2 * s.e\big(m'(x)\big)$ where $m'(x)$ denotes the slope of the regression function estimated by the local regression methods and s.e the standard error of the elasticity estimate. The standard errors in each year are estimated by bootstrapping (Efron and Tibshirani, 1993) with a modification that takes into consideration the clustered structure of the SUSENAS sampling procedure.[29]

One simple way of determining whether the elasticity estimates are significantly different at different levels of outlay is by checking whether the standard error bands for the 1996 estimate overlaps with standard error bands for the 1999 estimate. If at some range of outlay the standard error band for the 1999 estimate is clearly outside the standard error band for the 1996 estimate then it is safe to say that the elasticity estimate in 1999 is significantly different. The confidence interval bands around the estimated elasticity are wider at the right tails of the rural and urban distributions suggesting that the elasticity is measured with less precision at the very top ends of the distribution of per capita consumption. In both rural and urban areas, the income elasticity of the SSR in 1999 appears to be insignificantly different from that in 1996, especially for the households below the median level of per capita expenditures.

[29] For a detailed description of bootstrapping within a clustered sampling design see Deaton (1997) and Subramanian and Deaton (1996)

Figure A: Standard error bands around the income elasticity of SSR in 1996 and in 1999 in Rural and Urban Central Java, Indonesia





Annex B: First Stage Regressions

VARIABLES	(1) lnPCE	(2) lnPCE x D99	VARIABLES	(1) lnPCE	(2) lnPCE x D99
log (HH Asset Index)	0.24***	-0.02***	HH Head self employed with permanent assistance	0.20***	0.01
	(0.01)	(0.00)		(0.04)	(0.02)
log (HH Asset Index) x D99	-0.12***	0.14***	HH Head working without pay	0.00	0.00
	(0.01)	(0.01)		(0.02)	(0.01)
# Male children under 5	0.03	0.04	HH Head literate	0.00	0.04*
	(0.08)	(0.03)		(0.07)	(0.02)
# Female children under 5	-0.11	0.02	Spouse has no education	-0.04	0.01
	(0.08)	(0.03)		(0.03)	(0.01)
# Male children between 6-12	0.07	0.02	Spouse has not finished primary school	-0.09***	0.03***
	(0.08)	(0.03)		(0.03)	(0.01)
# Female children between 6-12	0.06	0.02	Spouse has completed primary school	-0.10***	0.02**
	(0.08)	(0.03)		(0.03)	(0.01)
# Male children between 13-19	0.26***	0.01	Spouse has completed junior/senior high school	-0.02	0.01
	(0.08)	(0.03)		(0.03)	(0.01)
# Female children between 13-19	0.27***	0.03	Spouse self employed without assistance	0.00	0.01
	(0.08)	(0.03)		(0.02)	(0.01)
# Males between 20-54	0.28***	-0.01	Spouse self employed with non-permanent assistance	0.04	0.02*
	(0.07)	(0.02)		(0.03)	(0.01)
# Females between 20-54	0.14**	0.01	Spouse self employed with permanent assistance	-0.07	-0.09**
	(0.06)	(0.02)		(0.06)	(0.05)
HH Head has no education	-0.20***	0.05**	Spouse working without pay	0.08**	0.03*
	(0.07)	(0.02)		(0.04)	(0.01)
HH Head has not finished primary school	-0.14***	-0.00	Spouse with wage employment	0.01	-0.00
	(0.03)	(0.01)		(0.02)	(0.01)
HH Head has completed primary school	-0.09***	0.01	Spouse literate	0.07**	-0.01
	(0.02)	(0.01)		(0.03)	(0.01)
HH Head as completed junior/senior high school	-0.06**	-0.01	Household size	-0.47***	-0.00

	(0.02)	(0.01)		(0.02)	(0.01)
HH Head self employed without assistance	-0.01	0.01	D99 (= 1 if year = 1999)	0.58***	12.52***
	(0.02)	(0.01)		(0.14)	(0.10)
HH Head self employed with non-permanent assistance	0.04*	0.01	R-squared	0.705	0.999
	(0.02)	(0.01)	F-statistic	148.3	24566

Note: The first stage regressions also include the control variables interacted with the 1999 year dummy. The coefficients for these interactions are not reported to conserve space.
*** p<0.01, ** p<0.05, * p<0.1