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An Empirical Test of Methods for Estimating Price Elasticities from Household Survey Data

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Abstract

Unit values, calculated as the ratio of household expenditure on a particular food to the quantity consumed, are often used as proxies for market prices in demand studies based on cross-sectional survey data. Because unit values are likely to give biased estimates of price elasticities, Deaton (1987, 1990) developed procedures for correcting these biases. However, empirical evidence on the bias created by unit values in demand systems is lacking. In this paper we use data collected specifically to carry out comparisons with the results of using market prices. Our findings suggest that unit values, whether used in naïve or improved estimation procedures, provide poor approximations to the elasticities calculated with market price data. We provide an example where optimal tax reform is distorted by using price elasticities calculated from unit values.

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I. Introduction

Applied researchers often use *unit values* from household surveys to estimate price elasticities of demand.[1] These unit values, which are calculated as household expenditures on a good divided by the quantity purchased, are used as proxies for market prices, especially in developing countries where market prices are either unavailable or have time series that are too short for estimating demand systems.[2] Ironically, price elasticities arguably are most needed in developing countries because of the relative importance of pricing policy which plays the same central role in fiscal policy that income tax and social security plays in most developed countries (Deaton, 1989). A matrix of price elasticities, which is needed to estimate the revenue effects of price reforms, can therefore provide socially useful information.[3]

Unit values, however, are not prices, and their use in demand studies is likely to create several biases (Deaton, 1987). In contrast to market prices, unit values reflect a household's quality choices, are affected by reporting errors and are unavailable for non-purchasing households. The quality effects matter because household surveys typically aggregate different varieties, so even if consumers faced the same prices, as the mix of varieties changes with variations in household income and other characteristics, the unit values change. In fact, unit values will tend to vary less than prices if consumers react to high prices by choosing lower quality, and this is likely to create a systematic overstatement in the absolute value of estimated price elasticities (Deaton, 1988). Reporting errors in either expenditures or quantities will also matter because they are reflected in unit values, and these errors are likely to cause spurious correlations with the demands that are being 'explained' by the unit values, causing estimated price elasticities to be biased.

Although methods for correcting the biases in demand elasticity estimates from unit value data have been developed, most notably by Deaton (1987, 1988, 1990) and more recently by Crawford et al. (2003), researchers rarely apply them.[4] Indeed, since Deaton's warnings about biases from unit values were published, more demand studies have used unit values without any correction than had been carried out previously. Some studies even use market prices for some goods and unit values for others in the *same* demand model, suggesting that researchers view the two as perfect substitutes (Minot and Goletti, 2000). One reason for this nonchalance may be that the magnitude of the bias from using unit values has never been empirically demonstrated. The biases that Deaton (1990) set out to correct were based on theoretical reasoning rather than empirical evidence. Similarly, the studies using Deaton's correction method have only had unit values available to them, so there has been no way to validate the approach. As Deaton (1990, p. 302) notes,

> "it would be extremely desirable to have data with direct measures of market prices against which this method could be compared."

In this paper we provide the first empirical evidence on such comparisons. Our results suggest that unit values provide poor approximations to the elasticities calculated from market prices, even when the correction method of Deaton (1990) is applied. These findings should be of relevance for those econometricians interested in demand systems analysis and those economists interested in agricultural commodity demand and tax reform in developing countries. The apparent failure of unit value-based methods suggests that new techniques (or new sources of data) for estimating price elasticities from household survey data are needed.

The data that we use were collected during a household survey in Papua New Guinea (PNG). This is an interesting country for conducting such comparisons because of the great price variation across space. PNG has no national road network. One-half of the population lives more than an hour from the nearest transport facility (World Bank, 1999). Recognizing the opportunity afforded by such a setting, and unlike most previous studies, we designed and managed the collection of the data ourselves, planning from the start to undertake a comparison of unit values with market prices. As such, the survey collected information on expenditures and quantities, as well as several different direct measures of price for PNG's main consumption commodities. We use the market price surveys as the standard against which the unit values are compared. This preference is not always apparent in the literature (Deaton and Grosh, 2000) but we believe that prices for well-defined items collected from market surveys using certain sampling rules are an appropriate standard. Moreover, three features of our case study country increase our faith in the market price surveys: villages are small and normally some distance apparent, and haggling is uncommon in markets in PNG. Both features mean that the prices observed by enumerators in the local market are likely to be the prices actually faced by households in the survey. Our demand data (from which our dependent variable is constructed) also were collected for a two week recall period during the same period of time during which the market price data were collected. As result, the fluctuation of prices over the season is not an issue.

The rest of the paper is organized as follows. Section II discusses the specification of the *base* demand model that uses measures of actual market prices to generate estimates of price elasticities. It then examines the different procedures that analysts apply to unit values, when actual prices are unavailable, paying particular attention to the method proposed by Deaton (1990). Section III describes the household survey data and focuses on explaining how we collected market prices and unit values. After examining the results and comparing the elasticities that come from models that depend on alternative procedures, Section IV compares the results of a tax reform analysis when alternative sets of elasticities are used. Section V extends the comparisons in several directions and also tests some of the key assumptions used by the Deaton correction methods. Conclusions are in Section VI.

II. Methods

The base model estimates price elasticities of demand using market prices and a "share-log" functional form (Deaton, 1989):

$$w_{Gi} = \alpha_{Gi} + \beta_G \ln x_i + \sum \theta_{GH} \ln p_H + \gamma' \mathbf{z}_i + u_{Gi} \tag{1}$$

where w_{Gi} is the share of the budget devoted to good G for household i, x_i is total expenditure of household i, p_H are the market prices (ones that are measured as actual prices rather than unit values) and $\mathbf{z}_i$ is a vector of other household characteristics. The choice of functional form has several considerations. First, we do not choose a log-log model even though it is analytical tractable and has the advantage that its parameters are elasticities which provide a simple, convenient and dimensionless way to measure price response. However, not all households consume all goods, and because we cannot take the logarithm of zero, the log-log model can only be used to describe the behaviour of those households who purchase positive amounts. For narrowly defined commodities, such as those used in this and many other household survey-based demand studies, such restrictions can eliminate a large fraction of the sample.

For certain applications (e.g., for estimating conditional demand elasticities), if one wanted to link this model to underlying utility theory using a data set with zero purchases, some researchers advocated using Tobit-type models (Heien and Wessells, 1990). Later work, unfortunately, showed that the Tobit approach leads to inconsistent estimates (Vermeulen, 2001).

In fact, Deaton (1997) argues that even with zero purchases by some households the link to utility theory is not necessary because all that is needed for the analysis of tax and subsidy reform is unconditional demand functions. Since the revenue effect of a tax increase does not depend on whether demand changes take place at the extensive or intensive margins, the analyst studying tax and revenue reform needs to include all households, whether they purchase or not (Deaton, 1990). Therefore, equation (1) is simply viewed as a linear approximation to the regression function of the budget share conditional on the right-hand-side variables, averaging over both zeros and nonzeros in much the same way that an aggregate demand function does (Deaton, 1997). The method of using budget share-based models and including households with zero purchases is also advocated and used by almost all other studies that have used Deaton's approach for estimating demand elasticities (e.g., Ayadi et al., 2003) and carrying out tax reform analysis (Nicita, 2004).

In this paper, our strategy is to initially estimate equation (1) and calculate *price elasticities*, ε_{GH}, using

$$\varepsilon_{GH} = (\theta_{GH} / w_G) - \delta_{GH}, \tag{2}$$

where δ_{GH} is the Kronecker delta (which is equal to 1 if G=H, and 0 otherwise) and budget share, w_G, is evaluated at its mean value. Likewise, the *expenditure elasticities*, ε_G, are calculated from:

$$\varepsilon_G = \beta_G / w_G + 1. \tag{3}$$

The base results using actual market prices are compared with the elasticities that result when the parameters of equation (1) are estimated by the following five unit value-based methods:

(i) using *unadjusted* unit values, on the subset of households recording consumption of each good, which we henceforth refer to as method UV1, an example of which is provided by Musgrove (1985);

(ii) replacing *missing* unit values with the *cluster mean* of the unit value (UV2--Sahn, 1988);

(iii) replacing *missing* unit values with the mean unit value calculated across other households in the same *region* and *season* (UV3--Minot, 1998);

(iv) replacing *missing* unit values with the *predictions* from a regression of observed unit values on regional and quarterly dummies and household total expenditures (UV4--Jensen and Manrique, 1998; Heien and Pompelli, 1989);[5]

(v) using *cluster mean unit values*, in place of both household-specific and missing unit values (UV5--Case, 1991; Rae, 1999).[6]

In addition to these methods, which replace unobserved prices in equation (1) with some form of unit value and then use the standard elasticity formula in equations (2) and (3), we also use the procedures developed in Deaton (1990). The Deaton procedure provides an alternative approach to estimating elasticities, the essence of which is captured by the following quotation:

> "Since household surveys typically collect data on *clusters* of households that live together in the same village and are surveyed at the same time, there should be no genuine variation in market prices within each cluster. Within-cluster variation in purchases and unit values can therefore be used to estimate the influence of incomes and household characteristics on quantities and qualities, and can do so without data on prices. Variation in unit values within the clusters can also tell us a good deal about the importance of measurement error. By contrast, variation in behaviour *between* clusters is at least partly due to cluster-to-cluster variation in prices, and this effect can be isolated by allowing for the quality effects and measurement errors that are estimated at the first, within-cluster stage."
> (Deaton, 1988, p. 419).

Unlike the traditional demand framework, Deaton's procedure starts with a two-equation system of budget shares (w_{Gic}) and unit values (v_{Gic}) that are both functions of the *unobserved* prices, (p_{Hc}):

$$w_{Gic} = \alpha^0_G = \beta^0_G \ln x_{ic} + \gamma^0_G \cdot z_{ic} + \sum_{H=1}^{N} \theta_{GH} \ln p_{Hc} + (f_{Gc} + u^0_{Gic}) \qquad (4)$$

$$\ln v_{Gic} = \alpha^{1}_{G} = \beta^{1}_{G} \ln x_{ic} + \gamma^{1}_{G} \cdot z_{ic} + \sum_{H=1}^{N} \psi_{GH} \ln p_{Hc} + u^{1}_{Gic} \qquad (5)$$

In addition to the variables previously defined, f_{Gc} is a cluster fixed-effect in the budget share for good G (this accounts for all cluster-specific factors that affect the budget share of the households within the cluster), u^{0}_{Gic} and u^{1}_{Gic} are idiosyncratic errors, and the c indexes clusters of surveyed households that live in close physical proximity to one another (such as households that live in the same village). Cluster fixed effects are not allowed in the unit value equation because they would obscure the link between unit values and the unobserved cluster prices.

There are two non-standard aspects of equations (4) and (5). First, the prices are unobserved. Second, consumers chose both quantity and quality, so that expenditure on good G is the product of price, quantity, *and* quality. Thus, if the logarithm of the budget share is differentiated with respect to $\ln x$ and $\ln p_H$ the results are not the usual expenditure and price elasticities, but rather:

$$\partial \ln w_G / \partial \ln x = \beta^{0}_{G} / w_G = \varepsilon_G + \beta^{1}_{G} - 1 \qquad (6a)$$

$$\partial \ln w_G / \partial \ln p_H = \theta_{GH} / w_G = \varepsilon_{GH} + \psi_{GH} \qquad (6b)$$

where ε_G is the elasticity of quantity demanded with respect to total expenditure (as in equation 3), ε_{GH} is the quantity elasticity with respect to the price of H (as in equation 2), β^{1}_{G} is the elasticity of the unit value with respect to total expenditure (henceforth, called the *quality elasticity*) and ψ_{GH} is the *elasticity of the unit value with respect to the price* of H.

Some intuition for the problems caused by the absence of price data is gained from noting that if equation (5) is re-written, switching $\ln v_{Gic}$ (from the LHS to RHS) with $\ln p_{Hc}$ (from the RHS to LHS), the coefficient on $\ln v_{Gic}$ would be ψ^{-1}_{GH}. Inserting this result into equation (4), if unit values are used in place of the unobserved prices in the budget share equation (as is done by most analysts), the coefficient on these unit values would not be θ_{GH}, but rather the ratio, $\psi_{GH}^{-1}\theta_{GH}$. The only case where the simple expedient of using unit values in place of prices gives accurate estimates of the log price derivative, θ_{GH}, is when the ψ matrix is an identity matrix (implying that prices and unit values move perfectly together).

Given these non-standard features, the implementation of Deaton's unit value-based method takes place in three stages. In the first stage, the procedure removes the household-specific effects of income and other demographic characteristics from the budget shares and unit values. To do so, equations (4) and (5) are estimated using OLS with dummy variables for each cluster in place of the unobserved price (a 'within' estimator). This specification controls for the cluster fixed effects, including those of unobserved prices, because the effect of market prices is not distinguishable from the effect of other cluster-varying effects in this model. Hence, the β_G^0, γ_G^0, β_G^1, and γ_G^1 parameters can be estimated consistently, even in the absence of market price data. These four sets of parameters to allow *adjusted* budget shares ($\hat{y}_{Gc}^0$) and unit values ($\hat{y}_{Gc}^1$) to be created with the quality effects due to income and other factors removed. These adjusted shares cannot yet be used, however, because they still contain measurement error (taken out in the second stage) and cluster-wide quality effects due to between cluster price variation (taken out in the third stage).

The first stage regressions also produce the residuals needed in the second stage for estimating the error covariances that are used to correct for the effect of any measurement error in unit values and budget shares. The error terms, e_{Gic}^0 and e_{Gic}^1, from equations (4) and (5) contain all the variability around the cluster means of w_{Gc} and v_{Gc} that is not explained by household characteristics, so this residual variability is assumed to reflect measurement error.

In the second stage, the adjusted budget shares ($\hat{y}^0_{Gc}$) and unit values ($\hat{y}^1_{Gc}$) are averaged *by cluster*:

$$\hat{y}^0_{Gc} = \frac{1}{n_c}\sum_{i \in c}\left(w_{Gic} - \tilde{b}^0_G \ln x_{ic} - \tilde{g}^0_G \cdot z_{ic}\right)$$
$$\hat{y}^1_{Gc} = \frac{1}{n^+_{Gc}}\sum_{i \in c}\left(\ln v_{Gic} - \tilde{b}^1_G \ln x_{ic} - \tilde{g}^1_G \cdot z_{ic}\right)$$

where n_c is the number of households in cluster c, n^+_{Gc} is the number reporting a unit value ($n_c \bullet n^+_{Gc}$ for all c) and tildes indicate estimates from the first stage. These cluster averages are then used to compute a 'between-cluster, errors-in-variables' regression:

$$\tilde{B} = \left(\tilde{S} - \tilde{\Omega}\tilde{N}^{-1}_{+}\right)^{-1}\left(\tilde{R} - \tilde{\Gamma}\tilde{N}^{-1}\right) \tag{7}$$

where the elements of $\tilde{S}$ and $\tilde{R}$ are the covariances of the cluster averages of the adjusted budget shares ($\hat{y}^0_{Gc}$) and unit values ($\hat{y}^1_{Gc}$), $\tilde{s}_{GH} = \text{cov}(\hat{y}^1_{Gc}, \hat{y}^1_{Hc})$ *and* $\tilde{r}_{GH} = \text{cov}(\hat{y}^1_{Gc}, \hat{y}^0_{Hc})$; $\tilde{\Omega}$ and $\tilde{\Gamma}$ are covariances of the errors from the first stage within-cluster residuals (e^0_{Gic}, e^1_{Gic}); and $\tilde{N}$ and $\tilde{N}_+$ are formed from the mean cluster size variables n_c and n^+_{Gc}.

After this second operation, Deaton's procedure has purged the unit values of the quality effects of income and demographics (the first stage) and measurement error (the second stage), but the adjusted shares and unit values are still contaminated by the influence of price on quality and are not yet able to produce the required parameters for calculating elasticities. This can be seen from Deaton's (1990) result that as the number of clusters, *c,* tends to infinity, holding fixed the number of households in each cluster, the probability limit of equation (7) is:

$$\underset{C \to \infty}{\text{plim}}\ \tilde{B} = \left(\Psi'\right)^{-1}\Theta' \tag{8}$$

In other words, at the limit, $\tilde{B},$ the estimated relationship from the second stage between-cluster regression of adjusted budget shares on adjusted unit values, is not the log price derivative, θ_{GH} (what we want) but rather the 'mixed' matrix $(\Psi')^{-1}\Theta'$.

The disentangling of the price and quality effects in equation (8) takes place at the third stage and relies on a separability assumption. This final step assumes that the effect of unobserved price on quality can be treated like an income effect:

$$\frac{\partial \ln v_{Gc}}{\partial \ln p_{Hc}} = \psi_{GH} = \delta_{GH} + \beta_G^1 \frac{\varepsilon_{GH}}{\varepsilon_G} \tag{9}$$

A price rise reduces the demand for a food group according to the price elasticity, ε_{GH}. When less is bought, the quality effect depends on the elasticity of quality with respect to expenditures on the group, given by the ratio of β_G^1 and ε_G. Substituting expressions for ε_{GH} and ε_G from equation (6) into equation (9) gives:

$$\psi_{GH} = \delta_{GH} + \frac{\beta_G^1(\theta_{GH}/w_G - \psi_{GH})}{(1-\beta_G^1) + \beta_G^0/w_G} \tag{10}$$

From equation (8) it is possible to express ψ_{GH} in terms of the elements of the parameter vector, *B*, and the matrix of log price derivatives, Θ. After using such a substitution to eliminate ψ_{GH} from both sides, equation (10) has a single unknown, θ_{GH}. This unknown is a function of the estimated parameter vector, *B*, the budget shares, *w*, and the first stage parameters β_G^0 and β_G^1.[7] Hence, the separability assumption in equation (9) allows the log price derivative, θ_{GH} to be identified, and this provides the remaining information needed to calculate the price elasticities.

III. Data

Data used in this paper come from the Papua New Guinea Household Survey (PNGHS), which was designed and supervised by the authors in 1995 and 1996. The key feature of this survey is that it collected information on *both* market prices and unit values for foods. The consumption block of the survey used a closed interval recall method; each household was interviewed twice so that the start of the consumption recall period was signalled by the first interview. The survey team collected consumption data on all food (36 categories) and other frequent expenses (20 categories). Respondents provided information on the value and quantity of food purchases, gifts, and own-production. Thus, in addition to the more usual *purchase unit values* (or the value of the goods purchased for cash in the market divided by the quantity of those goods), unit values for gifts and own-production are also available. The ratio of a household's value of consumption of a particular food to the quantity consumed takes account of all sources of acquisition and provides what we call a *consumption unit value*. In order to provide an estimate of household total expenditures, the survey also includes an annual recall of 31 categories of infrequent expenses and an inventory of durable assets, which provides estimates of the flow of services from durables.

Enumerators also collected market prices in each cluster using two different surveys. The prices of food items purchased from commercial retail outlets (e.g., rice, sugar, tinned fish, beer) and non-food items (e.g., soap, kerosene) were collected from the two main trade stores or supermarkets used by the households in each cluster. The survey instrument typically asked for prices for items based on finely defined specifications (e.g., a 340g can of "Ox and Palm" brand tinned meat). In some cases, however, if the only price available was for a non-specification good (e.g. a 200g can of Ox and Palm, or a 340g can of a different brand), the price of the closest substitute was recorded and we used these substitute prices to predict the missing price of the good of desired specification. The survey team also collected prices of locally produced and marketed foods from the nearest local outdoor market. Enumerators recorded the price and weight of up to six different lots of each commodity. The market price survey was carried out on two different days, providing us with prices on up to 12 lots of each food item for each cluster.

The survey covered a random sample of 1200 households, residing in 120 rural and urban clusters. Households in 20 of these clusters were re-interviewed approximately six months after the initial interviews and these are treated as separate clusters, giving a sample of 140 clusters. Unfortunately, we cannot work with clusters from the capital city where the survey differed in two important respects. First, the target number of households was obtained by taking fewer households per cluster (six, compared with 12 elsewhere). Second, in the capital city survey, one-half of the selected households in each cluster (3) were given diaries to record their expenditures, while enumerators visited the other half on the same two-visit, closed interval recall survey basis. Hence, for at least half of the sample (the diary-keepers), the unit values constructed from these records may have different measurement error characteristics than in the recall survey. Also, in a more densely populated urban area there is no guarantee that households buy in the market nearest their cluster (an assumption of our model). The marketing habits of many residents depend on where they work or go to school. Consequently, there is no reason to believe that all households in an urban cluster face the same prices. Given these sampling features, we work with the 96 clusters outside of the capital city.[8]

When selecting foods to include in the demand model, we had to make a critical choice about whether to take items directly from the survey on a fairly disaggregated basis or form broader aggregates of food types. Aggregating distinct items into a composite (e.g., root crops) highlights quality effects in unit values because inter-commodity variation is added to the existing intra-commodity quality differences. But combining distinct items interferes with the direct comparison of market survey prices with unit values because averaging across items would be involved and this may obscure measurement error. Since measurement error is noted by Deaton as likely to be the more important problem in practice, we choose to work with items directly from the survey. Also, if price elasticities are to be of practical assistance for reforming taxes and food subsidies, they need to be as specific as possible because otherwise there is risk of mis-targeting (e.g. subsidies leaking to the non-poor because a broader aggregate includes both luxuries and necessities).[9]

Although PNG residents consume a large number of crops across its diverse landscape, the study had to narrow its focus in order to construct a set of goods for which data are complete. A list of food items that includes all those that had an average share of at least 1.5 percent of the household's total consumption budget contains nine commodities (Table 1). But three of these foods (sago, yam and coconut) have market prices and consumption observed in only a few clusters. Another two foods (taro and betelnut) are consumed nationally, but we have market prices from only two-thirds of the clusters. These data constraints mean that the demand system used in the analysis is comprised of sweet potato, banana, rice, tinned fish, and "other goods", which is an aggregate of all other items in the survey.[10] The four individual foods comprise almost one-quarter of the average household's budget, and provide approximately one-half of the population's calories (Gibson, 2001). Although our purposes are mainly methodological, these four goods also have tax policy significance because until recently both rice and tinned fish were imported duty free, whereas other foods were subject to tariffs of at least ten percent. But following a switch to a Value-Added Tax (VAT), rice and tinned fish are now taxed at the same rate as other goods. The establishment of a local cannery also has caused tariffs on tinned fish to be set at a prohibitively high rate (70 percent--World Bank, 1999a).[11] In contrast, sweet potato and banana effectively fall outside of the tax net because they are almost all sold in the informal sector, where VAT is not collected. There are, however, proposals to use tariffs to dampen the demand for imported foods and stimulate the local production of these two staples (DAL, 2000) and these new issues raise interest in the cross-price elasticities.

Our data on these four food items are fairly complete; there are 80 clusters (containing 877 households) that have complete market price data and have at least one consumer per cluster for each of these four foods (Table 1). When the unit values from purchases are used instead of consumption unit values, the sample for the comparisons falls to 52 clusters. Correlations between market prices and consumption and purchase unit values demonstrate that the consumption unit value is no worse of a proxy for market price than is the more typically used purchase unit value (Table 2). Consequently, most of the comparisons use consumption unit values. Comparisons using purchase unit values will be presented as an extension in Section V.

The low correlation coefficients between market price and unit values suggest that either measurement error and/or quality effects (that are reflected in unit values) are important (Table 2). Although all correlations are significantly different from zero, they are considerably below 1.0. The correlations between price and purchase unit values range from 0.307 (tinned fish) to 0.537 (sweet potato), and for consumption unit values range from 0.310 to 0.454 (rows 1 and 2). Our data also demonstrate that the low correlations are not just the result of extreme outliers, so to the extent that there is measurement error, the error is subtle and potentially systemic.

Even after the original survey forms were re-examined and cases of data entry errors and obvious miscoding (e.g., kilograms entered as grams) were removed or rectified, we find that unit values appear to be noisy measures of cluster prices. For example, plots of price and unit values for tinned fish, the food item with the lowest level of correlation, show that the outliers are not driving the result (Figure 1). Extreme outliers are not evident even when comparing each household's unit value with the average across the *other* households in the cluster, which is the more feasible check in most surveys. To further ensure that the effect of measurement error on our regression results is not due to the effect of relatively small numbers of observations, in the next section we run a series of exercises to test how sensitive our results are to outliers. In our sensitivity analysis, we examine how deleting unit values more than two standard deviations from the mean affect our results, going beyond the spirit of Cox and Wohlgenant (1986) who only suggest deleting observations with unit values more than five standard deviations from the mean.

IV. Basic Results

Table 3 contains the budget share regression results when market prices are used (equation 1), along with the means and standard deviations of the explanatory variables. In addition to market prices and household expenditures, the regressions include household size (in log form), the share of the household in seven demographic groups: males and females 0-6 years, 7-14 years, 15-50 years, and over 50 years (males excluded), and dummy variables for the household head's gender (Female=1) and employment status (Work in the formal sector = 1). In Section V, we present the results (and draw comparisons with the results in this section) that also include dummy variables for the region in which the household is located and the quarter in which the household was surveyed.[12]

The elasticities derived from the regression results in Table 3 show well determined relationships between expenditures and demand and prices and demand (Table 4). Specifically, the own-price elasticities are all significantly different from zero (except for tinned fish where p=0.09—diagonal terms in columns 2 to 6). The own-price elasticities also are all significantly different than each other, except for banana and tinned fish where the hypothesis that the two elasticities are equal cannot be rejected (p=0.15). The lower precision of elasticities for tinned fish may be caused by the low variability in market prices for tinned fish (a feature of the commodity that also may contribute to the low correlation observed between its price and unit values--Table 2 and Figure 1). Several important cross-price elasticities, especially those for sweet potato, also are (column 2). The elasticity matrix in Table 4 also includes a final column and row for "other goods" based on the homogeneity and adding up restrictions (although homogeneity cannot be tested because of the lack of prices for the "other goods" category).[13]

In terms of policy implications, the elasticities in Table 4 can be divided into two groups: locally produced sweet potato and banana (henceforth, the *informally marketed* group), and the imported foods of rice and tinned fish (the *formally marketed* group).[14] Sweet potato and banana have low expenditure elasticities and are more likely to be consumed by the poor (the budget shares of sweet potato and bananas are 13.5 and 8.1 percent for those in the poorest expenditure quartile, more than twice their level in the richest quartile). In contrast, rice and tinned fish have higher expenditure elasticities and are consumed relatively more by those in the richest quartile.

Within either the informally or formally marketed group, however, the efficiency of taxing each individual food item is likely to differ. For example, the low own-price elasticity for tinned fish suggests that this is a more efficient good to tax than is rice. Alternatively, if the tax net could be widened to include the informally marketed foods, banana's low own-price elasticity would appear to make it a better candidate for taxing than sweet potato. These patterns would not be expected from more restrictive approaches, like the linear expenditure system, which enforces an approximate proportionality between expenditure elasticities and own-price elasticities.

The problem with using unit values for estimating elasticities becomes clear when examining the uniform shift to smaller price elasticities (in an absolute value sense) that results when we use unit values as proxies for market prices (Figure 2). The point estimates of the own-price elasticities for all six unit value procedures (including Deaton's) are biased towards zero.[15] The bias is especially apparent for sweet potato, where there is no overlap with the 95 percent confidence intervals surrounding the market price elasticity. Because quality variation is expected to make elasticities larger in absolute value (Deaton, 1988), the fact that the elasticities using unit values are attenuated (i.e., closer to zero) suggests that in the current setting the effect of bias due to measurement error is outweighing any quality effects.

Perhaps even more importantly from a policy perspective, when unit values are used, the ranking of food items, in terms of their own-price elasticities, changes (Figure 2). For example, banana is the most price inelastic under the naïve procedure of excluding observations with missing unit values (UV1). In contrast, sweet potato appears less price elastic than banana when cluster mean unit values (UV5) and the Deaton procedure are used.

The methodological insights from our results extend beyond highlighting the value of collecting actual market prices. They also show that the different methods of replacing missing unit values (UV2 to UV4) have less effect on elasticities than does the choice between using only non-missing unit values (UV1) or using cluster means (UV5—Figure 2). Surprisingly, the Deaton procedure provides no better point estimates of own-price elasticities than do the more naïve methods. One interpretation of this finding is that although theoretically elegant, in practice Deaton's method does not fix all problems with unit values. Alternatively, this result could be due in part to the small number of clusters in the PNG household survey.[16] In some sense, since Deaton's procedure reduces to a between-clusters regression, when there are relatively few clusters in the sample, the point estimates should be expected to be surrounded by wide standard errors (Deaton, 1990). If so, while providing an explanation for this particular sample, it would also imply that Deaton's correction procedure may have relatively few uses beyond those countries that have household budget data collected over a very large number of clusters.

For examining the performance of the alternative models using both own- and cross-price elasticities, there are too many cross-price elasticity estimates to display individually. Therefore, we calculate the aggregate bias (AB) as a summary indicator of the performance of each method.[17] To calculate the AB, we start by letting e be the vector of elasticities calculated from the market price data and $\hat{e}$ the corresponding elasticity vector from unit value data, with $AB = (\hat{e} - e)'(\hat{e} - e)$, which is just the sum of squared biases. We calculate the AB measure both for the own-price elasticities (AB1) and for the full system of own- and cross-price elasticities (AB2).[18] The aggregate bias in the own-price elasticities is greatest when estimation just uses the households with unit values available (AB1=2.454) and least when cluster means of the unit values are used (AB1=1.311—Table 5, column 1). When the cross-price elasticities are included in the aggregate bias calculation, the simple procedure of replacing missing unit values with cluster means does best (AB2=4.679), while the Deaton procedure does worst (AB2=10.275—column 2).[19]

Our above results suggest that measurement error in unit values is an important feature of our sample. However, removing unit values that are potential outliers does not appear to reduce the bias in the elasticities that result from the unit value procedures. The sample was increasingly "trimmed" by removing (log) unit values more than five, four, three and two standard deviations from the mean. The Deaton procedure was estimated on each of these trimmed samples. The aggregate bias for the own-price elasticities was 1.624 when (log) unit values more the five standard deviations from the mean were removed. When the threshold for trimming was reduced to four, three and then two standard deviations, AB1 was calculated as: 1.475, 1.835, and 4.002. Taking account of the cross-price elasticities, AB2 showed the same initial reduction and then increase with more severe trimming. Hence, the poor performance of the unit value methods does not appear to be due just to the presence of some outliers in the data.

Elasticity Estimation and Tax Reform Analysis

In this section, we use the estimates from the alternative approaches to examine how differences in estimated price elasticities would affect tax reform analysis. The theory of tax reform is concerned with small departures from an existing tax (or subsidy) regime. This theory requires a framework for judging the welfare effects of price changes due to taxes, and some accounting of the revenue effects of these tax changes. It is these revenue effects which depend on the response of consumers to price changes, as measured by the own- and cross-price elasticities. The costs (welfare effects) and benefits (revenue effects) of marginal tax changes can be summarised by l_i -- the social cost of raising one unit of government revenue by increasing the tax (or reducing the subsidy) on the *i*th good. If the ratio is large, social welfare would be improved by decreasing the price. To empirically estimate l_i, we follow Deaton (1997) and use:

$$\mathrm{l}_i = \frac{w_i^{\mathrm{e}} / \tilde{w}_i}{1 + \frac{\mathrm{t}_i}{1+\mathrm{t}_i}\left(\frac{\mathrm{q}_{ii}}{\tilde{w}_i} - 1\right) + \sum_{k \neq i} \frac{\mathrm{t}_k}{1+\mathrm{t}_k} \frac{\mathrm{q}_{ki}}{\tilde{w}_i}} \tag{11}$$

where t_i is the tax rate on good *i* (0.1 for rice and tinned fish, 0 for the others), q_{GH} is the log price derivative of the budget share (from equation (1) or the Deaton procedure), and the average budget shares w_i^{e} and $\tilde{w}_i$ are defined as:

$$w_i^{\mathrm{e}} = \left[\sum_{m=1}^{M} (x_m / n_m)^{-\mathrm{e}} x_m w_{im}\right] \Big/ \sum_{m=1}^{M} x_m \tag{12a}$$

$$\tilde{w}_i = \sum_{m=1}^{M} x_m w_{im} \Big/ \sum_{m=1}^{M} x_m \tag{12b}$$

where x_m and n_m are the total expenditure and size of household *m*, and e is the coefficient of inequality aversion. The expression for equation (11) is different from the more standard one (see, for example, Ahmad and Stern (1984), equation (38)) because it allows for both quantity and quality responses to tax-induced price changes.

According to the calculations in Table 5 (columns 3 to 6, row 1), when market prices are used to estimate q_{ki}, the highest ratio of social costs to benefits occurs when there is a marginal increase in the tax on rice (λ=1.58). Because the social cost of assessing a tax on rice is so high, tax reformers should consider either reducing the tax on rice or increasing taxes on other foods. But when the unit value procedures are used (rows 2 to 7), rice appears a more attractive candidate for higher taxes: Rice becomes the commodity with the second lowest λ-ratio (third lowest with the Deaton procedure). Hence, if unit values are used to calculate price elasticities, it would appear as if there is a relatively lower cost to taxing rice. Assuming the estimates in row 1 are correct, using unit values as proxies for market prices in an optimal tax reform exercise would lead policy makers in PNG to wrongly increase a tax which in fact is already a socially costly source of revenue.

V. Extensions and Further Tests

In order to examine the robustness of our findings, in this section we try a number of alternative approaches. If regional and quarterly dummy variables are added to the budget share regressions, most of the elasticities become smaller in absolute terms, while still preserving their initial ordering. For example, using market prices, the own-price elasticities become -1.28, -1.05, -1.55, and -0.96, for sweet potato, banana, rice and tinned fish (as compared with -1.59, -1.13, -2.20, and -0.61 when the dummy variables are excluded). Although the aggregate bias for most of the unit value procedures is lower once the regional and seasonal dummies are added (Table 6), the reduction is only in proportion to the reduced absolute value of the market price elasticities.[20] Moreover, the Deaton method appears to do considerably worse once the regional and seasonal controls are added, in part because it produces an estimated own-price elasticity for tinned fish of +1.63 (see appendix), as compared with the -0.96 estimated with market prices.

The other possible cause of poor performance for the unit value procedures is that they only work with purchase unit values rather than the consumption unit values used above. The results in the last four columns of Table 6, however, show this is not the case.[21] The comparisons in these tables rely on the restricted sample of 52 clusters with both market prices and at least one purchase unit value per cluster. For the six estimation methods there are 12 estimates of aggregate bias (AB1 and AB2 counted separately) and for five of these, the bias is larger with purchase unit values than with consumption unit values. Thus, the results in Section IV that used consumption unit values should not have been 'unfair' tests of the unit value methods.

Finally, in addition to allowing comparisons of the elasticities, the results using the market price data can also be used to test the separability assumptions of the Deaton method that are used to identify the unobserved price effects. Table 7 contains the results of testing the restrictions implied by equation (9), which relate the unobserved effects of market price on unit values to the observed income effects of quality and quantity. The first block of the table contains the empirically estimated elasticities of unit value with respect to prices, while the second block contains the value of this elasticity that is predicted by the right-hand side of equation 9 (i.e., using the quality elasticities, the price elasticities and the expenditure elasticities). The last block of the table contains tests of the discrepancies, which are significantly different from zero in five out of the 16 cases. Hence, this evidence does not appear to completely support the separability assumption used by Deaton (1988). This separability assumption, which is entailed in equation (9), is needed by the Deaton procedure to purge unit values of their quality effects, so evidence of its empirical relevance is useful for assessing the way that the Deaton deals with unobserved price effects.[22]

VI. Conclusions

This paper has presented evidence on the accuracy of price elasticities of demand that are estimated from household budget surveys, with unit values used as proxies for market prices. Such elasticities are increasingly being used as economists try to exploit one of the few data sources in developing countries that can help provide estimates of the demand responses to price changes that are needed for evaluating tax and subsidy reforms. Our findings suggest that unit values, whether used in naïve or improved estimation procedures, lead to biased estimates of the elasticities when compared to elasticities that would be calculated with actual market price data. We also have shown in the context of standard tax reform analysis, that the cost of having inaccurate measures of consumer price response is that the wrong commodity maybe taxed.

In one sense, our results may be doing nothing more than indicating small sample biases in Deaton's correction methods, which were already apparent from simulations (Deaton 1990, Figure 2). However, even knowing that a sample of 80 clusters is too small for Deaton's method to provide accurate estimates is a useful finding. Moreover, our results strengthen the case for conducting the sort of comparison reported here, but on a much larger sample of clusters, to fully evaluate the Deaton method.[23] The fairly high quality of our data (they were collected for this purpose) and the large, natural spatial variation in price between clusters in PNG, however, would argue that perhaps the poor performance of Deaton's approach is more than a small sample size problem. It could be that measurement error is a much more serious problem than the quality adjustments that are needed, and that Deaton's method is not able to correct fully for this combination of errors.

But because Deaton's approach is rarely used in practice, with economists typically settling for simpler procedures, our results should have a wider importance. These other methods of using unit values are not large sample methods and are often used with smaller samples than what is used here (e.g., Minot, 1998). Thus, our results suggest quite powerfully that naïve unit value procedures should be avoided if at all possible because they give biased estimates of price elasticities. The nonchalance of many applied researchers, in treating unit values as perfect substitutes for market prices, is unwarranted and more effort may be needed to development techniques for estimating price elasticities from household survey data.

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Figure 1: Comparison of Market Prices and Consumption Unit Values for Tinned Fish

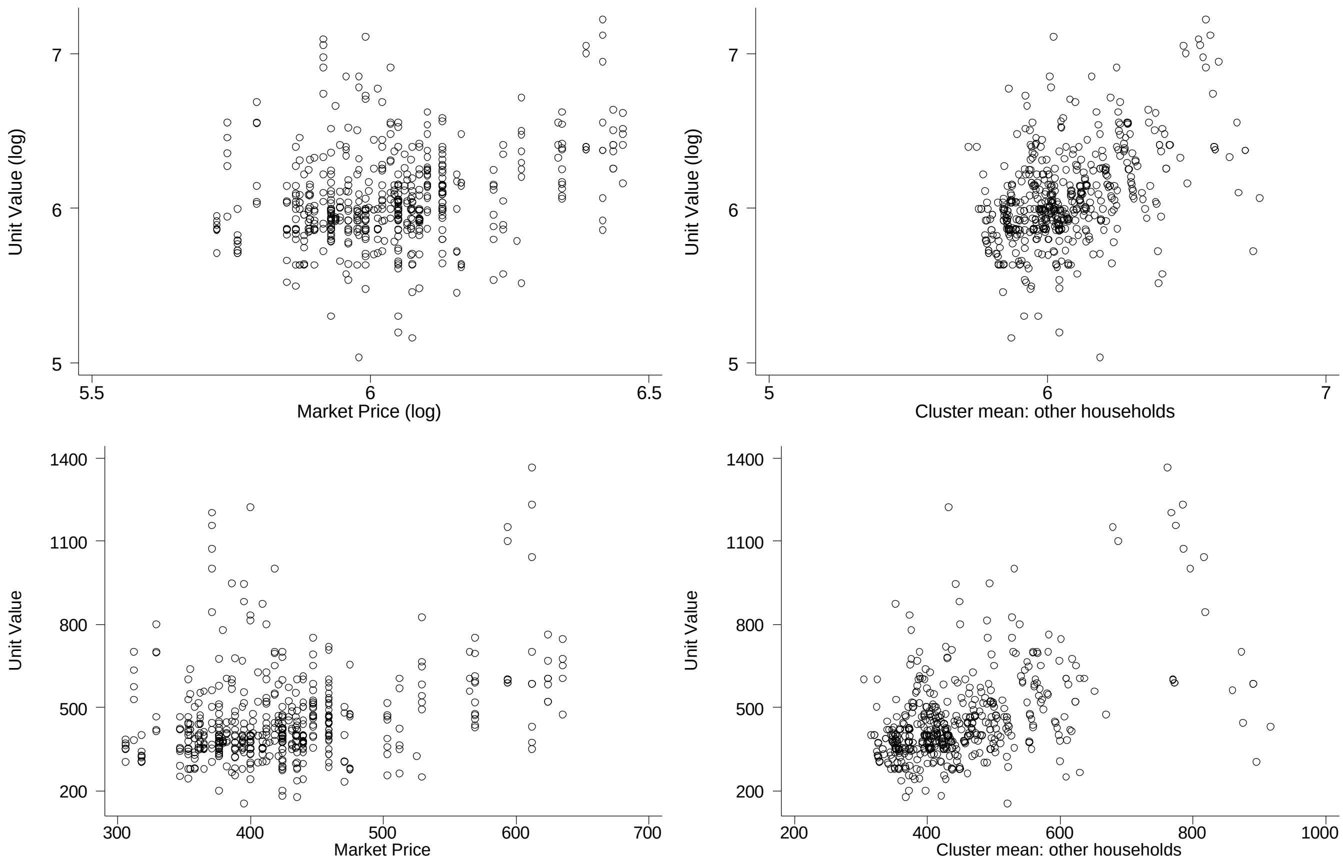

Figure 2: Comparisons of Own-Price Elasticities

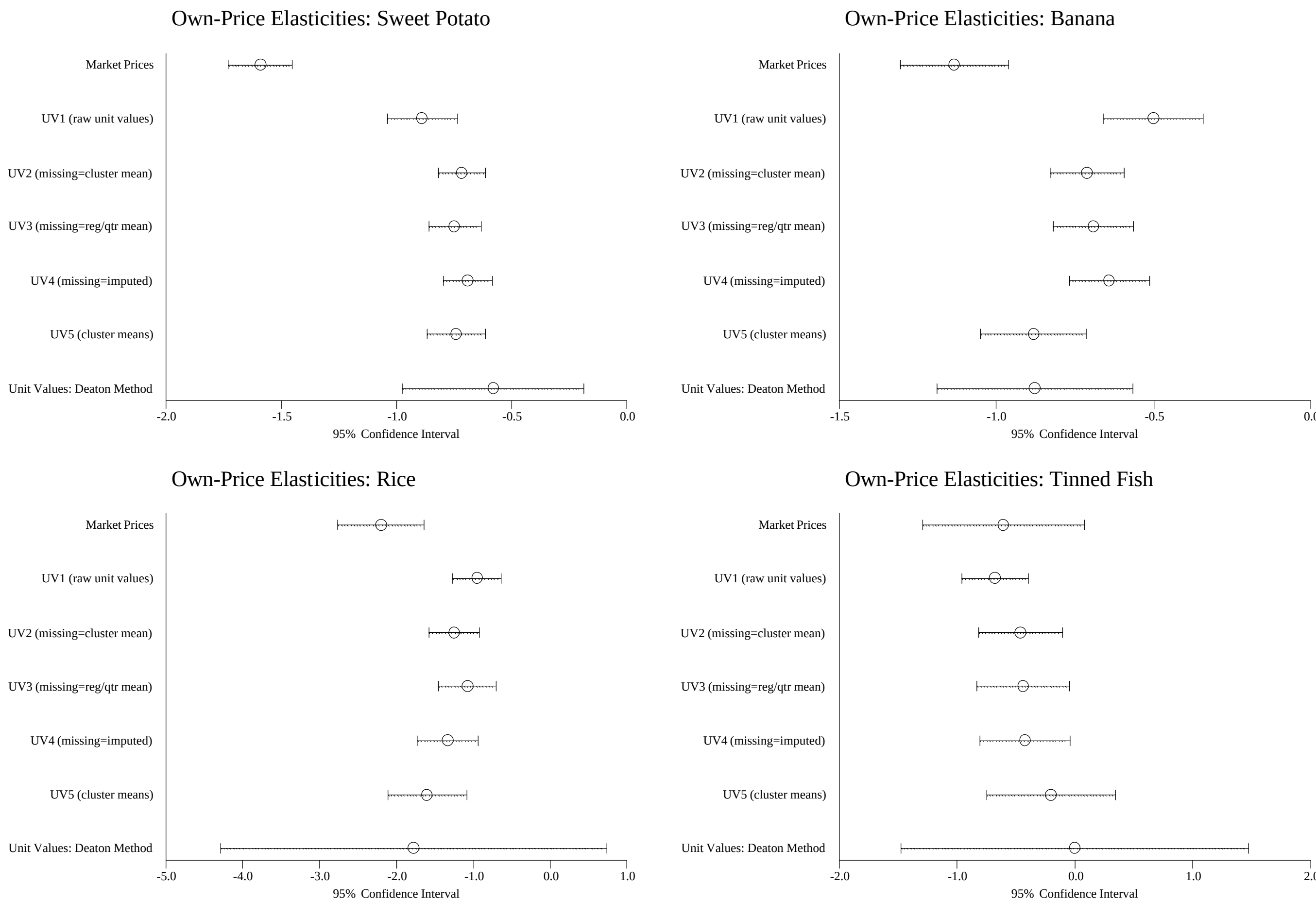

Table 1: Data Availability for Candidate Foods for the Demand System

	Average budget share	Number of clusters with:				
		Market prices observed (a)	≥ 1 household consuming (b)	Both (a) and (b)	≥ 1 household purchasing (c)	Both (a) and (c)
Sweet potato	**0.109**	**88**	**94**	**87**	**63**	**62**
Banana	**0.064**	**85**	**96**	**85**	**62**	**58**
Taro	0.050	62	88	60	36	25
Rice	**0.044**	**96**	**91**	**91**	**91**	**91**
Betelnut	0.034	66	94	65	83	60
Sago	0.022	24	41	23	21	14
Yam	0.020	26	59	22	11	9
Tinned fish	**0.017**	**94**	**91**	**89**	**90**	**88**
Coconut	0.015	42	78	39	49	27
Selected Items	0.235	84	89	80	56	52

Note: Selected items in bold.

Table 2: Pairwise Correlations Between Market Prices and Consumption and Purchase Unit Values.

Correlation pairs	Sweet potato	Banana	Rice	Tinned fish
Price, Purchase unit value	0.537	0.364	0.453	0.307
Price, Consumption unit value	0.454	0.405	0.423	0.310
Consumption unit value, Purchase unit value	0.697	0.712	0.916	0.977

Note: Prices and unit values are in logs. Based on the 80 clusters that have market prices and at least one consuming household for each of the four goods. (the maximum sample for any pairwise correlation is 877 households, but usually less because of missing unit values). All of the correlation coefficients are significant at the $p<0.001$ level.

Table 3: Data Description and Budget Share Regression Results using Market Prices, Papua New Guinea, 1996.

	Mean [std. dev]	Budget Share Regressions: Sweet potato	Banana	Rice	Tinned fish
ln price of:					
Sweet potato	3.131	-0.061	0.003	0.009	0.002
	[0.635]	(8.25)**	(0.69)	(2.30)*	(1.36)
Banana	3.428	0.094	-0.008	0.001	-0.002
	[0.511]	(10.18)**	(1.46)	(0.18)	(1.34)
Rice	4.785	0.161	-0.021	-0.058	-0.014
	[0.158]	(4.26)**	(1.26)	(4.14)**	(2.00)*
Tinned fish	6.040	-0.071	0.052	0.024	0.007
	[0.153]	(2.49)*	(3.05)**	(1.26)	(1.10)
ln total expenditure	8.005	-0.033	-0.020	-0.006	-0.002
	[0.884]	(5.87)**	(5.50)**	(1.82)+	(1.79)+
ln household size	1.617	0.029	0.006	0.009	0.003
	[0.552]	(3.10)**	(1.27)	(1.73)+	(1.49)
Share of the household who are:					
Male: 0-6	0.104	-0.014	-0.015	-0.037	-0.010
	[0.138]	(0.37)	(0.55)	(1.99)*	(1.08)
Male: 7-14	0.103	-0.029	-0.006	-0.014	0.006
	[0.129]	(0.60)	(0.20)	(0.62)	(0.57)
Male: 15-50	0.270	-0.028	-0.031	-0.015	0.008
	[0.191]	(0.80)	(1.31)	(0.79)	(0.97)
Female: 0-6	0.085	-0.077	-0.005	-0.030	-0.001
	[0.129]	(1.91)+	(0.16)	(1.44)	(0.08)
Female: 7-14	0.090	-0.068	-0.015	-0.021	0.007
	[0.122]	(1.53)	(0.50)	(0.94)	(0.68)
Female: 15-50	0.253	-0.032	-0.004	0.002	0.006
	[0.153]	(0.78)	(0.14)	(0.10)	(0.69)
Female: 51-	0.040	-0.006	-0.020	-0.033	0.003
	[0.119]	(0.14)	(0.56)	(1.26)	(0.22)
Characteristics of the household head					
Female	0.081	-0.037	-0.012	0.029	0.006
	[0.272]	(2.87)**	(1.11)	(2.96)**	(1.64)
Works in formal sector	0.260	-0.033	-0.010	0.010	0.006
	[0.439]	(4.12)**	(1.93)+	(1.76)+	(2.55)*
Constant		-0.108	0.036	0.192	0.050
		(0.18)	(0.31)	(1.82)+	(1.11)
R^2		0.26	0.10	0.06	0.03

Note: Absolute value of t-statistics in parentheses from heteroscedastically-consistent standard errors.
* significant at 5%; ** significant at 1%; + significant at 10%. N=877

Table 4. Expenditure and Unconstrained Price Elasticities using Market Prices, Papua New

Guinea, 1996.

	Expenditure Elasticities	Elasticity of quantity demanded with respect to the price of: Sweet Potato	Banana	Rice	Tinned Fish	Other Goods
Sweet Potato	0.68 (0.06)	-1.59 (0.07)	0.91 (0.09)	1.56 (0.37)	-0.68 (0.27)	-0.88 (0.34)
Banana	0.65 (0.06)	0.05 (0.07)	-1.13 (0.09)	-0.36 (0.29)	0.88 (0.29)	-0.09 (0.35)
Rice	0.88 (0.07)	0.18 (0.08)	0.02 (0.09)	-2.20 (0.29)	0.50 (0.39)	0.63 (0.36)
Tinned Fish	0.88 (0.07)	0.11 (0.08)	-0.14 (0.10)	-0.73 (0.37)	-0.61 (0.36)	0.49 (0.43)
Other Goods	1.08 (0.01)	-0.17 (0.04)	-0.03 (0.03)	0.03 (0.02)	0.01 (0.01)	-0.93 (0.06)

Note: Standard errors in parentheses. Results for "other goods" derived from homogeneity and adding up restrictions. Elasticities are evaluated at mean budget shares. Symmetry restrictions not imposed.

Table 5: Summary Comparisons of Estimates Using Market Prices and Unit Value Procedures

			Cost-benefit ratio (λ_i) of tax increase on:			
Data source and estimation method	AB1	AB2	Sweet potato	Banana	Rice	Tinned Fish
Market prices			1.32 (2)	1.41 (3)	1.58 (4)	1.10 (1)
UV1: Non-missing unit values	2.454	5.785	1.28 (3)	1.38 (4)	1.26 (2)	1.15 (1)
UV2: Cluster means if missing	1.865	4.679	1.35 (3)	1.38 (4)	1.26 (2)	1.05 (1)
UV3: Reg/qtr mean if missing	2.195	5.514	1.35 (3)	1.39 (4)	1.20 (2)	1.04 (1)
UV4: Impute if missing	1.840	4.938	1.35 (3)	1.39 (4)	1.25 (2)	1.04 (1)
UV5: Cluster mean unit values	1.311	5.731	1.36 (4)	1.36 (3)	1.33 (2)	0.96 (1)
Deaton method	1.624	10.275	1.39 (4)	1.34 (2)	1.36 (3)	0.88 (1)
Deaton method (applied to market prices)	0.017	0.061	1.31 (2)	1.42 (3)	1.60 (4)	1.11 (1)

Note: AB1 is the aggregate bias on the own-price elasticities, AB2 is the aggregate bias on own- and cross-price elasticities, λ_i is calculated from equation (8), using an inequality aversion parameter, ε=1. The values in ( ) are the good's rank in terms of λ_i, where "1" denotes the good with the lowest cost-benefit ratio from a marginal tax increase.

Table 6: Results From Extended Comparisons of Market Price Elasticities and Unit Value Methods

Data source and estimation method	With Region and Seasonal dummies		52 Cluster Sample			
			Purchase unit values		Consumption unit values	
	AB1	AB2	AB1	AB2	AB1	AB2
UV1: Non-missing unit values	1.566	3.847	2.814	8.763	2.545	4.906
UV2: Cluster means if missing	1.117	3.524	0.483	3.154	0.979	3.230
UV3: Reg/qtr mean if missing	1.577	4.517	0.355	3.473	1.147	3.529
UV4: Impute if missing	1.361	4.025	0.408	3.362	1.058	3.227
UV5: Cluster mean unit values	1.040	4.723	0.718	2.094	0.945	2.756
Deaton method	8.164	60.245	6.536	25.821	3.879	8.780

Note: AB1 is the aggregate bias on the own-price elasticities, AB2 is the aggregate bias on own- and cross-price elasticities.

Table 7: Test of the Separability Assumption Used to Identify Price Elasticities

Unit Value for:	Sweet Potato	Banana	Rice	Tinned Fish	
	Elasticity with respect to the price of:				
Sweet Potato	0.375	0.178	-0.934	0.396	
	(0.083)	(0.091)	(0.343)	(0.319)	
Banana	0.083	0.315	-0.380	-0.255	
	(0.072)	(0.099)	(0.268)	(0.318)	
Rice	-0.076	0.031	0.731	-0.052	
	(0.020)	(0.023)	(0.119)	(0.109)	
Tinned Fish	0.064	0.009	-0.131	0.577	
	(0.028)	(0.032)	(0.165)	(0.173)	
	Derived value from separability assumption				
Sweet Potato	0.576	-0.481	-0.679	0.725	
	(0.154)	(0.210)	(0.842)	(0.506)	
Banana	0.080	0.602	1.541	-1.307	
	(0.073)	(0.135)	(0.735)	(0.656)	
Rice	0.088	0.018	1.026	-0.182	
	(0.106)	(0.046)	(0.031)	(0.350)	
Tinned Fish	0.048	-0.039	-0.018	1.021	
	(0.078)	(0.067)	(0.001)	(0.025)	
	Wald test of discrepancy ($F_{1,79}$)				Joint test
Sweet Potato	2.24	8.88	0.06	0.20	5.21
	[0.14]	[0.00]	[0.80]	[0.66]	[0.00]
Banana	0.05	4.05	4.80	1.24	3.21
	[0.82]	[0.05]	[0.03]	[0.27]	[0.02]
Rice	2.24	0.04	12.09	0.25	3.28
	[0.14]	[0.85]	[0.00]	[0.62]	[0.02]
Tinned Fish	0.00	0.50	0.25	5.24	1.67
	[0.98]	[0.48]	[0.62]	[0.02]	[0.17]

Note: Standard errors in ( ) are consistent to heteroscedasticity and clustering. e_{GH} and e_G evaluated at the mean budget shares. *P*-values in [ ].

Appendix Table A. Expenditure and Symmetry-constrained Price Elasticities, Papua New Guinea, 1996

		Elasticity with respect to the price of:				
	Expenditure Elasticity	Sweet Potato	Banana	Rice	Tinned Fish	Other Goods
Sweet Potato	0.81 (0.05)	-1.51 (0.06)	0.21 (0.03)	0.11 (0.04)	0.01 (0.01)	0.37 (0.08)
Banana	0.67 (0.06)	0.37 (0.06)	-1.28 (0.09)	0.10 (0.07)	0.01 (0.03)	0.13 (0.12)
Rice	0.86 (0.07)	0.23 (0.08)	0.12 (0.09)	-1.69 (0.25)	-0.11 (0.14)	0.59 (0.30)
Tinned Fish	0.85 (0.07)	0.06 (0.08)	0.03 (0.10)	-0.30 (0.35)	-0.78 (0.35)	0.12 (0.41)
Other Goods	1.06 (0.01)	0.02 (0.01)	-0.02 (0.01)	0.03 (0.02)	-0.00 (0.01)	-1.09 (0.03)

Note: Standard errors in parentheses. Results for "other goods" derived from homogeneity and adding up restrictions. Elasticities are evaluated using parameter estimates and the mean budget shares from Table 3.

Endnotes

[1] There are too many studies to cite, but the idea of using unit values was first raised by Prais and Houthakker (1955), while Timmer and Alderman (1979) is the first application. There are now at least 40 published papers using unit values to estimate price elasticities and a list of these is available from the authors.

[2] Another response is to substitute assumptions, such as additivity, for data. For example, researchers often use the Linear Expenditure System to get price elasticities from household budgets, without any prices required. But additive preferences imply that expenditure and own-price elasticities are roughly proportional, forcing a tradeoff between equity and efficiency, and leading to recommendations of uniform rates of commodity taxes regardless of the patterns in the data (Deaton, 1997).

[3] The elasticities are not needed for evaluating the welfare effects of *marginal* tax and subsidy reforms. The existing demand structure, and some social weights for aggregating the effects across households, provides sufficient information when price changes are small (Ahmad and Stern, 1984).

[4] At least 45 published articles cite one or more of this set of papers, but other than Deaton and his co-authors, before 2000, the only published applications of the correction method appear to be Laraki (1990), Nelson (1994) and Gracia and Albisu (1998). Since 2000, we know of at least two published papers in English: Ayadi et al. (2003) and Crawford et al. (2003).

[5] A related procedure is to regress the deviation of household-specific unit values from the mean for each region in each quarter on a set of household characteristics and use this equation to predict adjusted unit values for the non-consuming households (Cox and Wohlgenant, 1986).

[6] A variant to minimise the influence of outliers is cluster medians (see Thomas and Strauss (1997) although at a broader geographical level and for disaggregated foods that are then aggregated into group price indices). A further variant is the cluster mean unit value for *other* households in the cluster as an instrumental variable correlated with the cluster price but not with the measurement error in each household-specific unit value. Benjamin (1992) uses this approach with farm wages, although it is restricted to clusters with at least two observed unit values.

[7] Equation (20) in Deaton (1990) provides the relevant formula in matrix notation.

[8] This sample may not give a powerful test of Deaton's procedure, which is consistent only for large numbers of clusters (see equation (6)). However, Deaton (1987) used these procedures for a sample from urban Côte d'Ivoire where the average number of clusters was only 86 (although subsequent applications have used hundreds or thousands of clusters).But there should be no lack of power for evaluating the other unit value procedures used in the literature, which do not appear to rely on large sample consistency claims and have been applied to smaller samples than the one here (e..g, Minot, 1998).

[9] For example, in urban Papua New Guinea, potatoes have an income elasticity of demand of 1.26 while for cassava it is only 0.22 (Gibson, 1998), yet both are likely to be grouped together in a 'root crops' aggregate.

[10] "Other goods" includes the five foods listed in Table 1 that do not have full information available on their prices. We are forced to assume that leisure is separable from goods demand because the survey did not gather data on wage rates.

[11] It is not clear whether this tariff has raised consumer prices or simply switched the source of supply from imports to the local cannery. While the nominal price of tinned fish rose 60% between September 1994 and March 1996, this appears to reflect the collapse in the PNG exchange rate rather than the establishment of the cannery. Between 1994 and 1996 there was no apparent rise in the price of tinned fish relative to the price of other canned foods.

[12] It is likely that some of the price differences between regions are the result of long-term influences (such as tastes and preference differences and agro-climatic factors that affect price and consumption levels), so adding these regional dummy variables could remove these long-run effects. The elasticities from the more comprehensive model are likely to be smaller in absolute value because they capture more of only the short-run effects and not the long-run ones (Deaton, 1997). Hence, a comparison of the results in Section IV and V may indicate whether unit values are more or less successful as proxies for market prices in the short-run versus the long-run.

[13] Symmetry restrictions can also be imposed to improve the precision of the estimates and the results of doing this are presented in Appendix Table A. The main effect of adding these restrictions is to attenuate several of the cross-price elasticities. Because the symmetry restrictions are rejected ($\chi^2_{(6)}$ =15.5) and given the focus is on the impact of using different data and not on the impact of the restrictions from demand theory, we do not use the symmetry-constrained elasticities in the comparisons.

[14] The local tin fish cannery uses mainly imported raw materials and fish.
[15] The only exception is for tinned-fish. Using unit values, with no correction for any of the problems (UV1), the calculated elasticity is slightly lower.
[16] When designing the PNGHS, we tried to increase the number of clusters to as great as possible. But, the high cost of moving among regions, and a limited budget, restricted the final cluster numbers. We had thought, however, that the higher quality of data and the large natural barriers among regions (which would provide more natural spatial heterogeneity) was enough to make up for the smaller number of clusters.
[17] See unpublished Appendix Tables A1-A7, which also include expenditure elasticities.
[18] In our AB calculation, we exclude the results for "other goods" which are simply derived from the other elasticities.
[19] To check that there was not some flaw in the programming, market prices were passed through the STATA code for the Deaton procedure. The results are summarised in the last row of Table 5 (the elasticity matrix is reported in full in unpublished appendix Table A7), and are very similar to the elasticities calculated from equation (1) and reported in Table 4, so there appears to be no obvious coding error causing the poor performance of the Deaton procedure.
[20] Specifically, the average *proportionate* error on the own-price elasticities is larger with the regional and seasonal dummies included than without them, for five of the six methods in Table 6. The elasticity matrices that provide the data for these comparisons are reported in unpublished Appendix B.
[21] The elasticity matrices that provide the data for this table are reported in unpublished Appendix C and D.
[22] Minten and Kyle (1999) also report evidence on changing relative prices within commodity groups that is inconsistent with the treatment of the quality effects in Deaton's model.
[23] Results are available from Vietnam, where 190 clusters are in the sample, and these show that the Deaton procedure and other methods of using unit values provide poor approximations to the elasticities calculated from market price data (Le, Gibson and Rozelle, 2002). However, the comparisons for Vietnam are clouded by the fact that the demands and unit values are based on an annual reference period, whereas the market prices may reflect seasonal conditions. Nevertheless, the similarity with our results, in a much more spatially integrated economy based on grains rather than root crops, add to the doubts about unit values that are raised by the results reported here.