Demand Systems With Unit Values:

Comparisons With Elasticities from Market Prices

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Abstract

The ratio of household expenditure on a particular food to the quantity consumed is often used as a proxy for market price in cross-sectional demand studies. These unit values are likely to give biased estimates of price elasticities, so Deaton (1990) developed procedures for correcting these biases. However, empirical evidence on the bias created by unit values in demand systems is lacking so in this paper we use data collected specially 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 actual market price data.

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I. Introduction

Income tax provides most government revenue in rich countries and transfer payments address equity concerns, so estimates of price elasticities of demand are not greatly needed for policy design. Poor countries, on the other hand, rely on indirect taxation and subsidies but the price elasticities needed to analyse the efficiency effects of these interventions are hard to obtain. The annual time series is too short – often 30 years at the most – to allow estimation of disaggregated demand systems. Hence, studies of tax reform in poor countries often use the linear expenditure system (LES) because it gives price elasticities from a single cross-section, using marginal budget shares and an estimate of committed expenditure, without any price data required (Ahmad and Stern, 1991). But the LES assumes additive preferences, which are not plausible for individual foods because they rule out inferior and complementary goods. Moreover, additive preferences imply that expenditure and own-price elasticities are roughly proportional (Deaton, 1974) 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).

To allow more flexible elasticity estimates for analysing tax and subsidy reform, attention has switched to other types of data, especially household budget surveys. These surveys typically provide estimates of expenditures *(E)* and quantities *(Q)*, at least for foods. The possibility of using unit values *(E/Q)* to estimate cross-sectional demand curves was first raised by Prais and Houthakker (1955, p.110). After a hiatus, Timmer and Alderman (1979) reactivated this line of research and since then, many studies have attempted to estimate price elasticities using such data.[1] These studies appear especially attractive in developing countries, where roads are bad and transport costs high,[2] causing far greater spatial price variation than in developed countries.

But unit values are not prices, despite the wishes of applied demand analysts that they might be, and several biases are likely to result from their use as proxies for market prices in demand studies. First, unit values reflect quality differences because household budget surveys typically aggregate different varieties, which may differ in quality. Even if consumers faced the same prices, as the mix of varieties changes across households, the unit value would change. Such variation does not reflect different conditions of supply, so it is undesirable for estimating demand curves (Cox and Wohlgenant, 1986). Indeed, unit values will tend to vary less than prices because consumers can react to high prices by choosing lower quality, leading to a systematic bias in estimated price elasticities (Deaton, 1988). Second, unit values are affected by reporting errors in either expenditures or quantitites and these errors are likely to cause spurious correlations with demand (Deaton, 1987). Finally, unit values are available only for purchasing (or consuming) households, requiring some procedure for imputing missing data.

Although methods for correcting the biases in demand elasticity estimates from unit value data have been developed, most prominently by Deaton (1987, 1988, 1990), practitioners rarely apply them.[3] One reason for this lack of interest by applied researchers may be that the magnitude of the bias from using unit values has never been empirically demonstrated. Despite the plea by Deaton (1990), there has never been a 'crucial experiment' in which the results calculated from market price data are compared with the results from either naïve or corrected unit value procedures.

In this paper we report on just such an experiment, using data that we have collected during a household survey in Papua New Guinea (PNG). This is an interesting country for conducting such comparisons because of the great spatial price variation, since PNG has no national road network and one-half of the population live more than an hour from the nearest transport facility (World Bank, 1999). Nature and government policy also have conspired to create a set of foods spanning the spectrum in terms of intra-commodity quality variation. At one extreme are foods covered by import or local production monopolies, such as rice, sugar, and tinned fish, which exhibit little quality variation. In contrast, other foods are quite diverse. For example, there are over 600 varieties of sweet potato. Finally, 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 measures of price for PNG's main consumption commodities.

The rest of the paper is organized as follows. Section II discusses the specification of our demand model and the different procedures that analysts use to overcome the problems of using unit values, paying particular attention to the methods proposed by Deaton (1990). Section III describes the household survey data, and focuses on explaining how we collected market prices and unit values. The basic estimation results and elasticity comparisons are in the following section. Section V extends the comparisons in several directions and also tests some of the 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):

$$w_i = \alpha_i + \beta_i \ln x + \sum \theta_{ij} \ln p_j + \boldsymbol{\gamma}'\mathbf{z} + u_i \qquad (1)$$

where w_i is the share of the budget devoted to good i, x is total expenditure, p_j are the prices (ones that are measured as actual prices rather than unit values) and **z** is a vector of other household characteristics. If one wanted to link this model to underlying utility theory, Tobit-type models of zero purchases might be needed (Heien and Wessells, 1990). But this link is not necessary because all that is needed for the analysis of tax and subsidy reform is unconditional demand functions; the revenue effect of a tax increase does not depend on whether demand changes take place at the extensive or intensive margins (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).

In this paper, our strategy is to initially estimate equation (1) and calculate price elasticities using

$$\varepsilon_{ij} = \left(\theta_{ij} / w_i\right) - \delta_{ij}, \qquad (2)$$

where δ_{ij} is the Kronecker delta (=1 if i=j, 0 otherwise) and budget shares are evaluated at their mean values. Likewise, the expenditure elasticities are calculated from:

$$\beta_i / w_i + 1. \qquad (3)$$

The base results are compared with the elasticities that result when the parameters of equation (1) are estimated by the following methods:

(i) using *unadjusted* unit values, on the subset of households recording consumption of each good (Musgrove, 1985);

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

(iii) replacing *missing* unit values with the mean unit value calculated across other households in the same *region* and *season* (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 (Jensen and Manrique, 1998; Heien and Pompelli, 1989);[4]

(v) using *cluster mean unit values*, in place of both household-specific and missing unit values (Case, 1991; Rae, 1999).[5]

In addition to these methods, which replace unobserved prices with some form of unit value and then proceed to use the standard elasticity formula, we also use the procedures developed in Deaton (1990). The Deaton procedure provides an alternative to equations (1), (2) and (3), and its essence 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).

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}):

$$w_{Gic} = \alpha_G^0 = \beta_G^0 \ln x_{ic} + \gamma_G^0 \cdot z_{ic} + \sum_{H=1}^{N} \theta_{GH} \ln p_{Hc} + (f_{Gc} + u_{Gic}^0) \qquad (4)$$

$$\ln v_{Gic} = \alpha_G^1 = \beta_G^1 \ln x_{ic} + \gamma_G^1 \cdot z_{ic} + \sum_{H=1}^{N} \psi_{GH} \ln p_{Hc} + u_{Gic}^1 \tag{5}$$

In addition to the variables previously defined, f_{Gc} is a cluster fixed-effect in the budget share for good G, u_{Gic}^0 and u_{Gic}^1 are idiosyncratic errors, and the i indexes households, the G and H index goods, and the c indexes clusters of surveyed households in close physical proximity. 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:

$$\partial \ln w_G / \partial \ln x = \beta_G^0 / w_G = \varepsilon_G + \beta_G^1 - 1 \tag{6a}$$

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

where ε_G is the elasticity of quantity demanded with respect to total expenditure, ε_{GH} is the elasticity of quantity demanded with respect to the price of H, β_G^1 is the elasticity of the unit value with respect to total expenditure (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 with $\ln p_{Hc}$ as the left-hand side variable, the coefficient on $\ln v_{Gic}$ would be ψ_{GH}^{-1}. Inserting this result into equation (4), if unit values are used in place of the unobserved prices in the budget share equation, 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 direct 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 calculations take place in three stages. First, equations (4) and (5) are estimated using OLS with dummy variables for each cluster (a 'within' estimator). This controls for the cluster fixed effects, and also for the unobserved prices because the effect of market price is not distinguished from the effect of other cluster-varying effects. 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. The residuals from these first stage regressions also provide the error terms, e_{Gic}^0 and e_{Gic}^1 needed in the second stage for estimating the covariances that are used to correct for the effect of any measurement errors in unit values and budget shares.

In the second stage, the effects of total expenditure and the demographics, *x* and *z*, are removed from the budget shares and unit values, using the parameter estimates from the first-stage. These adjusted budget shares and unit values are then averaged by cluster:

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

$$\hat{y}_{Gc}^1 = \frac{1}{n_{Gc}^+}\sum_{i\in c}\left(\ln v_{Gic} - \tilde{\beta}_G^1 \ln x_{ic} - \tilde{\gamma}_G^1 \cdot z_{ic}\right)$$

where n_c is the number of households in cluster *c*, n_{Gc}^+ is the number reporting a unit value and tildes indicate estimates from the first stage. These cluster averages are then used to compute a between-clusters, 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 and unit values; $\tilde{s}_{GH} = \mathrm{cov}(\hat{y}^{1}_{Gc}, \hat{y}^{1}_{Hc})$, $\tilde{r}_{GH} = \mathrm{cov}(\hat{y}^{1}_{Gc}, \hat{y}^{0}_{Hc})$; $\tilde{\Omega}$ 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}. The probability limit of equation (7) as the number of clusters, c tends to infinity, holding fixed the number of households in each cluster, is:

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

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 $\left(\Psi'\right)^{-1}\Theta'$. Even when unit values have been purged of the effects of income, demographics and measurement error, by using the first and second stages of Deaton's procedure, they are still contaminated by the influence of price on quality and so they do not yet provide the required parameters for calculating elasticities (Deaton, 1997).

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:

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

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} which 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.[6] 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 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. Consumption data were collected 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, 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.

Market prices were collected in each cluster using two different surveys. The prices of commercially produced food items (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 the cluster. These prices typically were for a finely defined specification (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 used to predict the missing price of the good of desired specification. The prices of locally produced foods were collected 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. Hence, the survey has prices on up to 12 lots of each food 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. But we do not 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, one-half of the selected households in each cluster were given diaries to record their expenditures, so any unit values constructed from these diaries would have different measurement error characteristics than in the recall survey. Also, in an urban area there is no guarantee that households buy in the market nearest their cluster, depending on where they may travel to work or school, and so there is no reason to believe that all households in the cluster face the same prices. Given these sampling features, we work with the 96 clusters outside of the capital city. This sample may not give a powerful test of Deaton's procedure, which is consistent only for large numbers of clusters (see equation (6)).[7] 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).

When selecting foods to include in the demand model we had to choose whether to take items directly from the survey 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 any measurement error. Since measurement error is noted by Deaton as 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).[8] Moreover, the trend in many household surveys, and especially the Living Standards Surveys, is to use a relatively short list of consumption items so future demand analysts may be more likely to work directly with the commodities that are defined in the surveys.

Table 1 reports details on major foods available from the survey. The food items in the table are all those that had an average share of at least 1.5 percent of the household's total consumption budget. The analysis drops three of the nine foods (sago, yam and coconut) because their market prices and consumption are observed in only a few clusters. Ecological constraints limit the production of these foods, and this has evidently limited the spread of their consumption. Although two other foods in Table 1, taro and betelnut, are consumed nationally, we have market prices from only two-thirds of the clusters. Including either taro or betelnut in a system of demand equations would substantially reduce the number of clusters and would weaken the power of the comparisons between elasticities from unit values and those from market prices.

Therefore, 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. The four individual foods comprise almost one-quarter of the average household's budget, and provide just under one-half of the calories (Gibson, 2001). Although our purposes are mainly methodological, these four goods also have 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 prohibitive rate of 70 percent (World Bank, 1999a).[9] In contrast, sweet potato and banana effectively fall outside of the tax net because they are 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.

Table 1 shows that there are 80 clusters (containing 877 households) with market prices and at least one consumer per cluster for each of these four foods. But if the unit values from purchases were used instead of consumption unit values, the sample for the comparisons would fall to 52 clusters. It is apparent from Table 2, which contains correlations between market prices and consumption and purchase unit values, that the consumption unit value is no worse of a proxy for market price than is the more typically used purchase unit value. Therefore, most of the comparisons will work with consumption unit values. Comparisons using purchase unit values will be presented as an extension in Section V.

The correlation coefficients in Table 2 are all considerably below 1.0, pointing to the likely errors and/or quality effects in the unit values. To ensure that these low correlations were not just the result of extreme outliers, 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. Even after these corrections, unit values appear to be noisy measures of cluster prices, as illustrated in Figure 1 for tinned fish, which is the food with the lowest correlations between prices and unit values. But there are not obvious outliers, 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. However, as a sensitivity analysis, the sample is trimmed, following Cox and Wohlgenant (1986) who delete 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 prices and household expenditures, the regressions include (log) household size, 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 whether the household head was either female or employed in the formal sector. In Section V, results will be presented that also include dummy variables for the region the household is located in and the quarter in which they were surveyed. It is likely that some of the price differences between regions are the result of long-term influences, so adding these regional dummy variables could remove these long-run effects. The resulting elasticities are likely to be smaller in absolute value because they refer to the short-run (Deaton, 1997), so a comparison of the results in Section IV and V may indicate whether unit values are any more successful as proxies for market prices in the short-run than in the long-run.

The expenditure, own- and cross-price elasticities of quantity demand that are derived from the Table 3 regression results and the mean budget shares are reported in Table 4a. The own-price effects are well determined (except for tinned fish) as are several cross-price effects, especially for sweet potato. The lower precision of the own-price elasticity for tinned fish may be caused by the low variability in market prices for tinned fish. From the descriptive statistics in Table 3 it can be inferred that the coefficient of variation of prices is lowest for tinned fish and highest for sweet potato. The elasticity matrix in Table 4a also includes a final column and row for "other goods" based on the homogeneity and adding up restrictions (homogeneity cannot be tested because of the lack of prices for the "other goods" category). Symmetry restrictions can also be imposed to improve the precision of the estimates and the results of doing this are presented in Table 4b. 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.

In terms of policy implications, the elasticities in Table 4 can be divided into two groups: locally produced sweet potato and banana, and the imported foods of rice and tinned fish.[10] Sweet potato and banana have low expenditure elasticities and are more likely to be consumed by the poor; rice and tinned fish have higher expenditure elasticities. But within these two groups, the efficiency of taxing each food is likely to differ. The low own-price elasticity for tinned fish suggests that this is a more efficient good to tax than is rice. If the tax net could be widened to include the informally marketed foods, banana would appear to be a better candidate than sweet potato.[11]

These patterns would not be expected from the linear expenditure system, which enforces an approximate proportionality between expenditure elasticities and own-price elasticities.

How do the results compare when unit values are used? The point estimates and 95 percent confidence intervals reported in Figure 2 suggest that all six unit value procedures cause an upward bias (towards zero) in the own-price elasticities, especially for sweet potato .[12] Because quality variation is likely to make elasticities larger in absolute value (Deaton, 1988),[13] the fact that the elasticities using unit values are attenuated (i.e., closer to zero) suggests that in the current setting, the effect of measurement error is outweighing any offsetting biases due to quality effects. The ranking of foods is also changed, with banana appearing the most price inelastic under the naïve procedure of excluding observations with missing unit values, while sweet potato appears less price elastic than banana when cluster mean unit values and the Deaton procedure are used. The different methods of replacing missing unit values show less impact than the choice between using only non-missing unit values or using cluster means. Also, considerable imprecision is apparent in the estimates from the Deaton procedure, which essentially reduces to a between-clusters regression. With relatively few clusters in the sample it is not surprising that the point estimates are surrounded by wide standard errors.

There are too many cross-price elasticity estimates to display individually, so the aggregate bias (AB) is calculated as a summary indicator of the performance of each method.[14] Let $\boldsymbol{\varepsilon}$ be the vector of elasticities calculated from the market price data and $\hat{\boldsymbol{\varepsilon}}$ the corresponding elasticity vector from unit value data, so that the bias is $\hat{\boldsymbol{\varepsilon}} - \boldsymbol{\varepsilon}$, and $AB = (\hat{\boldsymbol{\varepsilon}} - \boldsymbol{\varepsilon})'(\hat{\boldsymbol{\varepsilon}} - \boldsymbol{\varepsilon})$, which is the sum of squared biases. The aggregate bias is calculated both for the own-price elasticities (AB1) and for the full system of own- and cross-price elasticities (AB2), but excluding the results for “other goods” which are simply derived from the other elasticities. According to the results in Table 5, 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). 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).[15]

Removing unit values that are potential outliers does not appear to reduce the bias in the unit value procedures. The sample was increasingly severely trimmed by removing (log) unit values more than five, four, three or 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.

The biased elasticity estimates when unit values are used as proxies for market prices are likely to distort optimal tax reforms. The last four columns of Table 5 contain estimates of the social cost-benefit ratios, λ_i of a marginal increase in tax on each of the four foods, calculated from:

$$\lambda_i = \frac{w_i^{\varepsilon} / \widetilde{w}_i}{1 + \frac{\tau_i}{1+\tau_i}\left(\frac{\theta_{ii}}{\widetilde{w}_i} - 1\right) + \sum_{k \neq i} \frac{\tau_k}{1+\tau_k} \frac{\theta_{ki}}{\widetilde{w}_i}} \tag{11}$$

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

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

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

where x_m and n_m are the total expenditure and size of household *m*, and ε is the coefficient of inequality aversion.[16] According to the calculations in Table 5, when market prices are used to estimate θ_{ki}, the highest ratio of social costs to benefits occurs when there is a marginal increase in the tax on rice (λ=1.58). But when the unit value procedures are used, rice appears a more attractive candidate for higher taxes with the second lowest λ-ratio (third lowest with the Deaton procedure). Hence, 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 is already a socially costly source of revenue.

V. Extensions and Further Tests

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.[17] 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, 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. But the results in the last four columns of Table 6 show this is not the case.[18] 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.

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 relates 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). 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.[19]

VI. Conclusions

This paper has presented evidence on the accuracy of price elasticities of demand 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 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 that would be calculated with actual market price data.

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.

But because Deaton's approach is rarely used in practice, with economists typically setting 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 which may lead to erroneous analysis of tax and subsidy reform in poor countries.

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

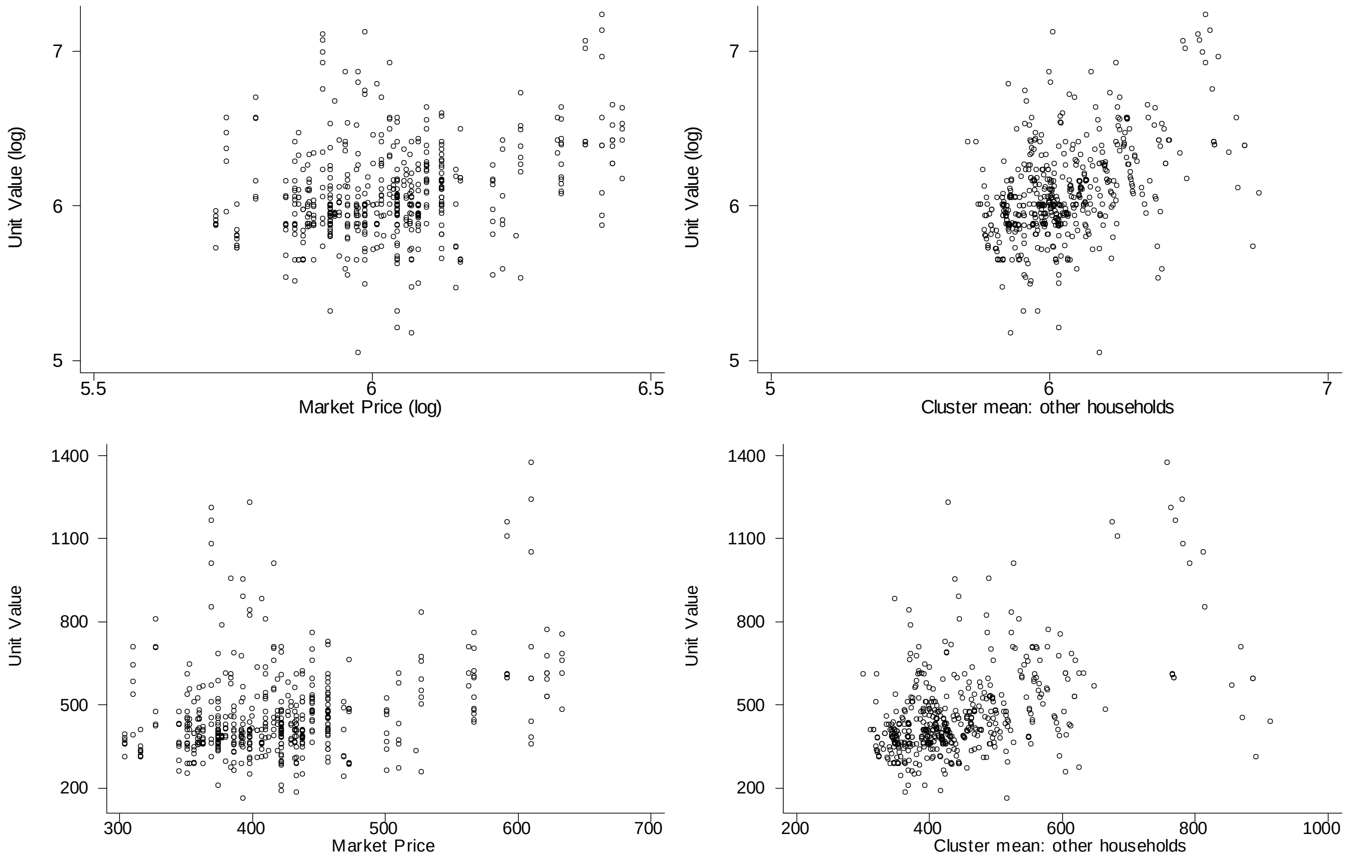

Figure 2: Comparisons of Own-Price Elasticities

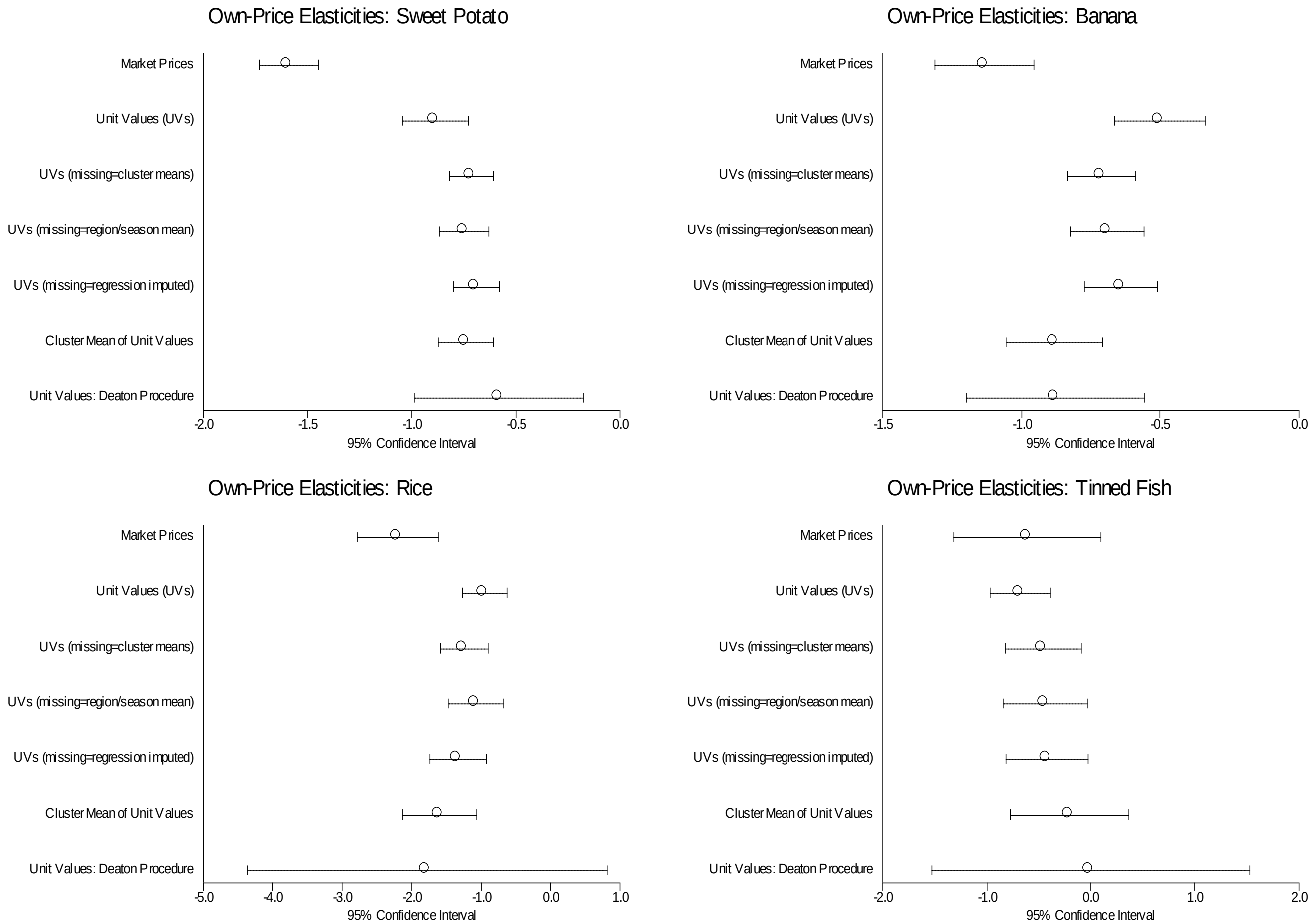

Table 1: Data Availability for Candidate Foods for the Demand System

		Number of clusters with:				
	Average budget share	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

	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 –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

	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 4a. Expenditure and Unconstrained Price Elasticities, Papua New Guinea, 1996

	Expenditure Elasticity	Elasticity 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 ( ). Results for "other goods" derived from homogeneity and adding up restrictions. Elasticities are evaluated at the mean budget shares.

Table 4b. Expenditure and Symmetry-constrained Price Elasticities, Papua New Guinea, 1996

	Expenditure Elasticity	Elasticity with respect to the price of: 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 ( ). Results for "other goods" derived from homogeneity and adding up restrictions. Elasticities are evaluated at the mean budget shares.

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

Data source and estimation method	AB1	AB2	Cost-benefit ratio (λ_i) of tax increase on: Sweet potato	Banana	Rice	Tinned Fish
Market prices			1.32 (2)	1.41 (3)	1.58 (4)	1.10 (1)
Non-missing unit values	2.454	5.785	1.28 (3)	1.38 (4)	1.26 (2)	1.15 (1)
Cluster means if missing	1.865	4.679	1.35 (3)	1.38 (4)	1.26 (2)	1.05 (1)
Reg/qtr mean if missing	2.195	5.514	1.35 (3)	1.39 (4)	1.20 (2)	1.04 (1)
Impute if missing	1.840	4.938	1.35 (3)	1.39 (4)	1.25 (2)	1.04 (1)
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 - 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		52 Cluster Sample: Consumption unit values	
	AB1	AB2	AB1	AB2	AB1	AB2
Non-missing unit values	1.566	3.847	2.814	8.763	2.545	4.906
Cluster means if missing	1.117	3.524	0.483	3.154	0.979	3.230
Reg/qtr mean if missing	1.577	4.517	0.355	3.473	1.147	3.529
Impute if missing	1.361	4.025	0.408	3.362	1.058	3.227
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. ε_{GH} and ε_{G} evaluated at the mean budget shares. *P*-values in [ ].

Notes

[1] A single journal, the *American Journal of Agricultural Economics*, has at least 20 such papers.
[2] In the former Zaire, transport costs make up one-quarter to one-third of wholesale prices (Minten and Kyle, 1999).
[3] At least 45 published articles cite one or more of this set of papers, but other than Deaton and his co-authors, the only applications of the correction method appear to be Laraki (1990), Nelson (1994) and Gracia and Albisu (1998).
[4] 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).
[5] 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.
[6] Equation (20) in Deaton (1990) provides the relevant formula in matrix notation.
[7] 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).
[8] 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.
[9] 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.
[10] The local tin fish cannery uses mainly imported raw materials and fish.
[11]Alternatively, sweet potato appears to be the better candidate for shifting the supply curve down, in terms of the allocation of agricultural research budgets.
[12] The only exception is for tinned-fish. Using unit values, with no correction for any of the problems, the calculated elasticity is lower.
[13] This can be seen from equation (4) and (5). If unit values are used in place of the unobserved prices in the budget share equation, the coefficient on these unit values would not be θ_{GH}, but rather the ratio, $\psi_{GH}^{-1}\theta_{GH}$. Because consumers can downgrade quality in response to price increases, it is expected that $\psi_{GH} < 1$, so the estimated coefficient when using unit values is larger than when using actual prices.
[14] See Appendix Tables A1-A7, which also include expenditure elasticities.
[15] To check that there was not some flaw in the programming, market prices were passed through the STATA code for the Deaton procedure. The elasticities (reported in full in Table A7) are very similar to the elasticities calculated from equation (1) and reported in Table 4, so there appears to be no obvious error in the code which is causing the poor performance of the Deaton procedure.
[16] This expression for the cost-benefit ratio of a marginal tax increase is adapted by Deaton (1997) from the more usual one (see, for example, Ahmad and Stern (1984), equation (38)) and allows for both quantity and quality responses to tax-induced price changes.
[17] 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 Appendix B.
[18] The elasticity matrices that provide the data for this table are reported in Appendix C and D.
[19] 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.