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# Estimating the relationship between food prices and food consumption – methods matter

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## ABSTRACT

Concerns about the growing prevalence of obesity worldwide have led researchers and policy makers to investigate the potential health impact of fiscal policies, such as taxes on unhealthy foods. A common instrument to measure the relationship between food prices and food consumption is the price elasticity of demand. Using meta-regression analysis we assessed how differences in methodological approaches to estimating demand affected food price elasticities. Most methodological differences had a statistically significant impact on elasticity estimates which stresses the importance of using meta-estimates or testing the sensitivity of simulation outcomes to a range of elasticity parameters before drawing policy conclusions.

## Introduction

Food prices and consumers' responses to changing food prices have gained substantial attention in recent years, particularly in the context of introducing fiscal policies to tackle unhealthy diets associated with rising prevalence of obesity and non-communicable disease globally. (Basu et al. 2014, Briggs et al. 2013, Leifert and Lucina 2015, Manyema et al. 2014, NiMhurchu et al. 2015, Tiffin and Arnoult 2011, Zhen et al. 2014) These policies can include both taxes on unhealthy foods or beverages, and subsidies on healthy alternatives. Also, the potential effect of "carbon" taxes on foods, the production of which is associated with high levels of greenhouse gas emissions, is another area of growing interest where consumers' responses to relative price changes through taxes, is studied (Briggs et al. 2016, Green et al. 2015, Säll and Gren 2015, Wirsenius, Hedenus, and Mohlin 2011). To evaluate the effectiveness of this type of policies it is crucial to know the extent to which consumers change consumption patterns as a response to changes in prices.

The key instrument to predict consumer response to food price changes is the set of own- and cross-price elasticities (OPE's and CPE's). Both OPEs and CPE's are needed to estimate the impact of price changes on consumption patterns which later feed into simulation models. The own-price effect, which in the policy context, is the direct intended impact of a tax or a subsidy, is generally larger in comparison to cross-price effects. However, cross-price effects are equally important as these can reinforce the own-price effect (i.e. complement or budget effect) or work in the opposite direction (i.e. substitute effect). If substantial and significant, these less predictable indirect effects can affect policy implications of the simulation outcomes (Cornelsen et al. 2014). As an example, our previous work found that in high-income countries a 10% increase in the price of sweets (including sugar-sweetened beverages) was associated with a reduction in its consumption by 5.6% but a 3% increase in consumption of cereal, dairy and fruits and vegetables, off-setting nearly half of the calories lost from reduced sweets consumption (Cornelsen et al. 2014). In contrast, in low-income countries, a similar price increase for sweets was associated with a 7.4% reduction in its consumption and an increase in the consumption of other foods by 6.1%. As the share of sweets in providing daily calories is much lower in low-income countries (7% in comparison to 13% in high-income countries), the substitution towards other foods, in particular cereals, far exceeded the reduction in calories from lower sweets consumption. If considering calorie intake as an outcome, the case for taxing sweets in high-income countries becomes much weaker, considering that nearly half of the calories are substituted to other sources. However, in low-income countries where under-nutrition is of concern, an increase in the price of sweets has an unexpected effect of increasing the total calories via substitution to relatively cheaper and staple foods.

In order to use price elasticities when simulating policy effects, researchers have to either use previously published estimates or estimate these from available data. While numerous studies exist estimating the demand for foods and beverages aggregated into broad groups, there is a lack of good quality evidence on specific and detailed food items, such as sugar-sweetened beverages, or products with high sugar, fat and salt content. This problem is aggravated in low-income countries where also source data are less available. For aggregate food groups, for which more estimates are available, the researchers still face a difficult choice in choosing between models using different source data, taking different underlying assumptions, and thus applying varied methods and functional forms. In such cases, using meta-estimates combining the findings from available studies could provide more robust estimates. Equally, when estimating elasticities from food expenditure or other consumption data, researchers face similar challenges in choosing the most appropriate data and methods from available alternatives.

The wide range of such alternatives, differing levels of complexity in methods and reports on known sources of bias in demand system estimations (Deaton 1988, Cox and Wohlgenant 1986, Shonkwiler and Yen 1999) have led us to question if, and to what extent, there exist systematic differences in the estimated food price elasticity values depending on the methods applied. Few previous studies have attempted to analyse this using the meta-regression approach. Gallet (2009, 2010) analysed variations in the OPEs of meat (Gallet 2010) and fish (Gallet 2009) demand. Chen et al. (2015) analysed both OPEs and CPEs of demand in China for 12 aggregate food groups, alcoholic beverages and tobacco (Chen et al. 2015). All three studies used slightly different explanatory variables in the meta-regression but found significant effects on elasticity estimates from variables describing data type and structure, model structure, model specification, estimation methods and publication type.

In our previous work we conducted a systematic review of literature estimating the demand for foods and beverages and provided meta-estimates for OPEs and CPEs for aggregated food groups in low-, middle- and high-income countries (Green et al. 2013, Cornelsen et al. 2014). In this study we employed the same global database of food price elasticities, extending over 12 years, to investigate and discuss in detail the influence of various methodological aspects on the estimates of both OPEs and CPEs using meta-regression analysis.

It has to be noted that it is particularly important to focus on the impact of the difference in methodological approach on CPE estimates. Changes in own prices have a more noticeable impact on consumption while the marginal impact of price change of a single alternative good is harder to capture. Also, CPEs found in the literature show a high degree of heterogeneity, including switches from positive (substitute goods), to negative (complementary goods). Hence, the bias can potentially cause a change in the direction of the elasticity, but this will be difficult to detect because the sign of the cross-price elasticity cannot be assumed *a priori* for most foods.

## Methodology

We used OPE and CPE estimates from a database of food price elasticities compiled from a systematic literature review conducted with an end date in August 2011 for OPEs and in November 2012 for CPEs (both data sets are available upon request from authors) [9,10]. Searches for studies in the review were done in academic databases (ISI Web of Science, EconLit, Medline, AgEcon and Agricola) and in other online resources (Google (and Scholar), Ideas, Eldis, websites of USDA, FAO, World Bank and IFPRI).

The review included published and grey literature, with English abstracts, estimating food price elasticities of demand using data from 1990 onwards and applying multiple equation methods. It included studies that used nationally representative aggregate data (national average statistics), data from household surveys (cross-sectional) or data from longitudinal surveys. It is important to note that as the criteria prescribed the inclusion of studies employing only post 1990 data, a number of studies employing long time series data, dating back in cases to 1950's, were excluded. While this ignores historic literature, it avoids any systematic differences in elasticities across a long period of time due to vastly changed economic conditions that affect the relationship between food prices and purchasing decisions.

A further distinction in estimated elasticities is between uncompensated (Marshallian) and compensated (Hicksian) elasticities. The latter is of interest when the focus is specifically on price effects net of the income effects. Because of their direct policy relevance, we used only the uncompensated, Marshallian elasticities that combine both price and budget effects.
The uncompensated (Marshallian) own- and cross-price elasticities were extracted and aggregated into nine broad categories of food – fruits and vegetables; meat; fish; cereals; dairy; eggs; fats and oils; sweets, confectionery and sweetened beverages (sweets); and other

foods. Price elasticities for food groups at a higher aggregation level than that used in this study (e.g. ‘meat and dairy’) and cross-price elasticities that, due to aggregation, were within one food group (e.g. cross-price elasticity of pork to beef price) were excluded. Price elasticities that were reported across different sub-population groups were averaged.

The database included also the following information on the included studies: whether the study was published in a peer-reviewed journal, country and region of the study, data source and type and years, function and estimation type in the demand analysis and whether the demand system estimated was complete or conditional. Countries were assigned into low-, middle- and high-income countries following the classification by (Muhammad et al. 2011).

For the purposes of this study additional, more detailed information on data and methods applied in the same set of studies were extracted: data frequency, whether and how censoring in the data was controlled for, which type of data were used for prices, and whether potential biases were addressed in the price data.

*Methodological aspects of demand analysis*

There are numerous methods available to estimate the demand for consumer goods and the choice largely depends on the theoretical and empirical assumptions the researchers are willing to make, and on data availability. The systematic review described above, and thus this paper, focused on research employing multiple equation methods for demand analysis, in coherence with current economic theory on consumer behaviour, prescribing that consumers allocate their fixed budget across the available bundle of goods depending on relative prices. Thus, demand functions for different goods are not independent from each other, and demand for a specific good is influenced by the price of all goods. This requires the joint estimation of demand equations as errors are correlated and cross-equation constraints exist. These demand systems can range from a subset of particular foods or beverages (e.g. different meats or beverages) or they can include the whole range of consumer goods, where the former type reflects 'conditional' demand and the latter relates to complete demand.

In the analysis we considered following known sources of bias as well as other aspects that may exert a systematic influence on price elasticity estimates:

*Different data structures*

The structure of data used to estimate demand systems varies from aggregate time series of national food expenditure data to very detailed consumer data recorded with hand-held scanners for all purchases of sample households. The level of detail in the data can have an effect on the estimated elasticties as cross-sectional data are unable to capture the dynamic components of consumption while time series data can suffer from aggregation bias (Denton and Mountain 2001, Blundell, Pashardes, and Weber 1993). We considered three types of data structure a) aggregate (national average statistics including time series), b) household survey data (cross-sectional) and c) longitudinal survey data (panel). As in individual studies data are often manipulated (e.g. aggregated), we also tested whether the frequency of the time dimension had an impact on the elasticity estimates using three categories of monthly or more frequently, quarterly and annual.

*Functional form*

Different functional forms for estimating demand systems can lead to different elasticity estimates (Dameus et al. 2002). The most popular demand systems stem from the Almost Ideal Demand System (AIDS). The AIDS model is non-linear in prices, but linear in total expenditure and most studies adopt a linearized version (LA-AIDS) due to its simple implementation (Deaton and Muellbauer 1980), although this linearization has been also associated with potential biases in certain situations (Pashardes 1993). In more recent years the quadratic version (QAIDS) has become popular, as it allows for a non-linear relationship between income and expenditure across different income groups (Banks, Blundell, and Lewbel 1997). However, other systems are also used, often to address theoretical considerations or specific data issues. For example, the translog model is similar to AIDS but requires a larger data set as the number of parameters to estimate is higher (Barten 1993, Deaton 1986), whereas the LinQuad incomplete demand system is more flexible and imposes fewer restrictions on theoretical consumer preferences in comparison to AIDS (Pan, Mohanty, and Welch 2008). Mixed Demand models assume that for some products the prices are given but for some others it is the quantity that is given and prices adjust to clear the market (e.g. suitable for quickly perishable foods) (Moschini and Rizzi 2005). Endogeneity of quantities, prices and budget can also be accommodated in dynamic demand systems estimated through time series econometric techniques such as cointegrated demand systems (Pesaran and Shin 2002).

*Estimation method*

Different estimation methods may also determine elasticity estimates. Because of correlated errors, demand systems are typically estimated via seemingly unrelated regression (SUR), or full information maximum likelihood (FIML). However, some studies address dynamics, habit formation and/or price and/or income expenditure endogeneity by adopting instrumental variable methods, such as two-stage least squares (2SLS) or – more recently –the aforementioned cointegrated demand systems (VEC-AIDS).

*Conditionality of the elasticities*

Complete demand systems may be estimated in a single stage, or can be broken down into two or more subsequent stages of budget allocation. For example, Edgerton (Edgerton 1997) assumed a three-step budgeting decision where in the first step the decisions are made on how much is spent on foods compared to non-food items (health, housing etc). In the second step the budget for foods is divided into major categories (e.g. fruits) and in the third step the budget is allocated between individual expenditure to individual food items (e.g. orange juice). Elasticities that are estimated from a single-stage complete system are unconditional (i.e. price changes of individual food items affect decisions of expenditure on all consumer goods) whereas elasticities that are estimated from demand systems only at second or third level are conditional on the expenditure at higher level (i.e. price changes affect decisions on expenditure within the food group).

Edgerton (Edgerton 1997) reported that restricting the analysis to the last stage of the multi-stage budgeting process can lead to considerable errors, and suggested correction procedures which are rarely adopted. Rickertsen (Rickertsen 1998) and Klonaris and Hallam (Klonaris and Hallam 2003) both report deviations between conditional and unconditional elasticities indicating possible systematic differences.

*Censored data*

If demand systems are estimated using household level data, it is likely that the dataset is censored (i.e. non-expenditure is observed). This can be due to genuine and deliberate non-consumption driven by preferences and independent from prices and incomes (e.g. vegetarianism), non-consumption during the survey period (especially for low-frequency consumptions and/or short survey period) or non-consumption explained by price and income level (i.e. at a different price/income level consumption would occur). Including these zero-observations without corrections has been shown to lead to biased estimates of the price elasticities (Heien and Wessells 1990). The most common approach to address the bias is to estimate the demand in two steps (Shonkwiler and Yen 1999) where the first step is the dichotomous decision on whether to consume or not and in the second stage the decision on how much to consume is taken, or to include a correction term in the demand equations, based on a Heckman-type correction procedure (Heien and Wessells 1990).

*Use of unit values as a proxy for price data*

As price data are often missing, particularly in household surveys, unit values, calculated as a ratio of expenditure to its quantity is a common type of price indicator used. This approach offers a solution to missing price data and provides variability in prices that using aggregate consumer or retail prices at one point in time (e.g. cross-sectional data) may not provide (Deaton 1988). Unit prices also mean that there are no discrepancies between the price and consumption data (Deaton and Grosh 2000). However, unit values are affected by quality bias and may lead to inconsistent estimates because errors in unit values are correlated with errors in the expenditure share or quantity data also employed in the model (Deaton 1988). Quality bias can arise because the goods purchased are generally at least to some extent aggregated (e.g. beef rather than specific cuts) and households at higher income levels might be purchasing more expensive (higher quality) beef cuts compared to poorer households. Any price change is likely to affect both decisions on quantity and quality of the foods.

The approaches to adjust for this bias assume that households in the same geographical area and at the same point in time face the same prices. A basic adjustment is based on regressing unit values on household socio-demographic characteristics to disentangle the quality, quantity and price effects (Cox and Wohlgenant 1986), while a more theoretically consistent approach requires the joint estimation of quantity and quality demand functions (Deaton 1988). Because consumers respond to price changes by adjusting their quality allocation, the price variation captured by unit values is usually smaller than the actual one. This means that any consumption response is ascribed to a downward biased estimate of price change, hence generating an overestimate of elasticities.

*Meta-regression model*

To explore the influence of these methodological approaches separately for OPEs and CPEs we estimated two meta-regression models. To account for study level heterogeneity we estimated a two-level random intercept model where the individual elasticities represented the second level, and study, the first level. The model was fitted using maximum likelihood (ML) with bootstrapped standard errors (50 replications). The dependent variable was the uncompensated OPE or CPE. Independent variables that were used in the model, describing the methodological approaches, are summarised in table 1.

Multicollinearity across the independent variables was tested for using the variance inflation factor (VIF). Variables with VIF values above 10 in the model were removed through testing various model specifications. The best model was chosen based on the highest value for adjusted coefficient of determination ($R^2$) and lowest vales for VIF.

Extreme values of elasticities, defined as lying outside of the absolute value of three standard deviations of the mean, within the food group, were considered as outliers. This led to a removal 1.7% (n=47) and 2.41% (n=131) of the observations from OPE and CPE datasets, respectively.

**Results**

The final database included 130 studies estimating OPEs (n=2,749) and 78 studies reporting CPEs (n=5,191) for any of the nine food groups.  The electronic supplement describes each included study in more detail. Table 1 shows the distribution of the variables within the dataset. A large share of OPEs (66%, n=1,803) were from two multi-country studies using International Comparison Program Data (IPCD)(Muhammad et al. 2011, Seale, Regmi, and Bernstein 2003) while CPEs the two largest studies counted only for 28% of observations.

*Table 1 here*

For both OPEs and CPEs, there were more estimates from grey literature, largely conference papers. OPEs were more often estimated for low-income countries while more CPE estimates were available from high-income countries. This is likely due to more detailed data being available from high income countries allowing for more detailed food items to be included. Approximately one third of both OPE and CPE estimates were from Europe.

When the two ICPD studies, estimating unconditional elasticities, were excluded, elasticities were most commonly estimated from complete models (CPE) or conditional on food sub-group expenditure (OPE). Household survey data (cross-sectional) was the most common data structure and annual data frequency was most common for both types of elasticities, even if the ICPD studies were excluded. The majority of elasticities were estimated with a version of the AIDS function if excluding the ICPD studies where the Working Preference Independence (Florida) model was employed. The most common estimation type was SUR if the two big studies were not considered and ML if these were included (OPEs only).

Two-step methods were the most common approach to deal with censored data. For 8% of OPEs (31 studies) and 18% of CPEs (23 studies) it was not reported whether censoring was dealt with (or if it was an issue) but based on the structure of the data used was a possible problem. Also, 46% of OPEs (64 studies) and 40% of CPEs (40 studies) were estimated using unadjusted unit values as approximations for price data, or price data had not been described at all. Lastly, both OPEs and CPEs were mostly estimated for fruits and vegetables or meat and the average data year used in estimation of elasticities was 2000 for OPE's and 2001 for CPE's, respectively.

*Meta-regression results: own-price elasticities*

Table 2 presents the meta-regression results for OPEs. The Likelihood Ratio (LR) test indicated that study level effects were statistically significant ($p<0.001$) justifying the use of a two-level model. We excluded the variable describing data type as it was leading to multicollinearity in the model and data frequency alone yielded a higher value for adjusted $R^2$ in comparison to data type. Since OPEs entered the model with their original (negative) sign, a positive coefficient indicates a lower elasticity (i.e. less sensitive demand to changes in prices) and a negative coefficient indicates a higher elasticity (i.e. more sensitive demand to changes in prices).

*Table 2 here*

As expected, OPEs indicated less sensitive demand to price changes as country income level increased with an average difference of 0.27 between the food price elasticity in low-income countries and high-income countries ($p<0.001$). In comparison to Europe, OPEs from Africa and Asia indicated more sensitive food demand to changes in prices. Differences between Europe and Australasia, North- or South-America were not significant at conventional levels.

Both monthly and quarterly data were associated with higher OPEs (i.e. more sensitive demand to changes in prices) in comparison to annual data ($p<0.05$). Choice of estimation type was jointly significant ($p=0.011$) in explaining some of the variation in elasticity estimates although individually only the ‘other estimation method’ was significantly different (higher elasticity) in comparison to elasticities estimated using SUR method ($p=0.001$). To the contrary, the type of price data was jointly not significant at conventional levels ($p=0.279$) although we found OPE estimates from retail price data to be less elastic ($p=0.015$). This is confirmative evidence that using unadjusted unit prices, as a proxy for retail prices, leads to an overestimation of OPEs in comparison to using actual retail price data.

OPE estimates were also affected by whether or not censoring in the data was addressed. In comparison to two-step methods, aggregating data or using any other method was associated with less elastic OPEs ($p<0.001$). Equally, when it was not reported how censoring was addressed or where it was not applicable (e.g. aggregate data), the elasticities were associated with less elastic values ($p<0.001$).

Factors that were not associated with significant changes (at the 5% level) in elasticity estimates were whether the study was peer reviewed, whether elasticities were conditional or unconditional, function type employed and mean year of data.

*Meta-regression results: cross-price elasticities*

As the sign of CPE is not predictable, meaning that there is no theoretical prior on whether foods are complements or substitutes, and the estimates are generally much smaller compared to own-price elasticity estimates, the interpretation of the meta-regression results presented in table 3, is more complicated and cannot be compared to the *a priori* expectations. Similarly to the OPE model, multicollinearity was detected in the model leading to exclusion of variables describing data type and country income level. Study level effects were equally found to be significant ($p<0.001$).

CPEs from peer-reviewed studies were weakly associated with more positive values in comparison to grey literature ($p=0.063$). Regional differences were also detected for CPEs. In comparison to Europe the CPEs were more positive in Asia ($p<0.001$), North-America ($p=0.013$) and South-America ($p=0.004$).

*Table 3 here*

Monthly or more frequent data were associated with more positive CPE values ($p=0.012$) in comparison to annual data, but no significant differences were detected between quarterly or annual data. LS estimations were associated with smaller elasticities in comparison to models estimated by SUR (p=0.017). However, jointly, the estimation type was significant only at the 10% level.

Similarly to the OPEs, the way of addressing censoring in consumption data was found to jointly explain part of the variation in CPEs (p<0.001). At the individual level, only studies where censoring was not applicable (e.g. employing aggregate data) were associated with smaller cross-price elasticities (p<0.001).

The type of price data used also explained part of the variation in CPEs (p<0.001). Adjusted unit prices were associated with more positive cross-price elasticities (p<0.001) in comparison to unadjusted unit prices. The coefficient for retail price was also positive but not significant at conventional levels (p=0.291). Studies applying other price data (see section 3 for details) were associated with more negative CPE estimates (p=0.007). Mean year of data, function type and the conditionality of elasticities, equally to OPEs, were not associated with changes in elasticity estimates at conventional statistical significance levels.

**Discussion**

There are many individual studies estimating the price sensitivity of food demand across the globe.  Only a few have attempted to synthesise this body of research (Andreyeva, Long, and Brownell 2010, Cabrera Escobar et al. 2013, Chen et al. 2015, Cornelsen et al. 2014, Gallet 2010, 2009, Green et al. 2013) and all these analyses have pointed to the wide array of data and methods used in the estimation of price elasticities, which inevitably leads to a question how this affects the sensitivity of the elasticity estimates, particularly when used in policy simulations.

We have added to the literature by using a meta-regression analysis and a large existing data base to examine how methodological differences affect OPE and CPE estimates after controlling for food group, study specific effects, country income level and study region, and whether studies were peer-reviewed. While individual studies in economics have explored the bias in demand analysis of different methodological aspects, the meta-regression analysis approach allowed us to combine these and to explore the influence on the elasticity estimates in a single model.

Similarly to the few previous studies using the same approach (Gallet 2010, 2009, Chen et al. 2015), we found that the different methodological approaches to a smaller or larger extent do matter as these significantly affect food price elasticity estimates. We found statistically significant differences in OPEs estimated using data at different frequencies and estimated by different estimation methods. The latter was also found to be an important influence in the previous two meta-regression analyses of OPEs (for fish and meat only) (Gallet 2010, 2009) and in the analysis of Chinese food price elasticities (Chen et al. 2015).

The method of addressing censoring in the data, led to significant differences in OPE estimates. In particular, using a two-step demand system was associated with smaller (more sensitive) OPEs in comparison to aggregation of data or where no adjustments were done. This finding has relevant implications for future studies as increasingly more disaggregated data is collected and analysed, such as scanner data, which by its nature is highly censored.

For both OPEs and CPEs the type of price data used was associated with significant differences. As the theory predicts, quality adjusted unit values and retail prices led to larger (less sensitive) OPE estimates in comparison to using unadjusted unit values. Hence, attention should be given to which price data are used and whether adjustments for quality differences need to be implemented.

Interestingly, we did not find evidence of significant influence stemming from the choice of functional form or conditionality of the elasticities. However, the functional form was defined only by two categories because the types of models that were non-AIDS were relatively few as by selection criteria only studies using a demand system were included. Similarly, to Chen et al. we found that published papers had significantly more positive CPE's which may indicate some publication bias and certain expectations to the estimated values.

In comparison to OPEs, the impact of methodological bias on CPEs can be more serious as CPEs can switch from negative to positive with a different interpretation for either case (substitute or complement products). CPEs are usually considerably smaller (not far from zero) and thus even small bias can cause the switch in the direction of the effect that in the worst case can lead to a different policy suggestion. This particularly affects studies modelling the potential impact of health- or environment-related food taxes or subsidies where it is necessary to explicitly include cross-price effects to understand the changes across the whole diet, rather than just taxed or subsidised products. If the demand estimation provides inconclusive CPE estimates or estimates that are close to zero, simulation studies should test the sensitivity of their findings by allowing both negative and positive cross-price effects to test the bounds of the outcome measures. Alternatively, meta-estimates, such as provided by (Green et al. 2013, Cornelsen et al. 2014, Gallet 2010, 2009, Andreyeva, Long, and Brownell 2010, Chen et al. 2015, Cabrera Escobar et al. 2013, Clements and Si 2015) should be used.

**Concluding Comments**

We conclude that studies wishing to employ food price elasticities as parameters in their simulation or other exercises should be careful in choosing these from previous literature or in the choice of methods to be used in the estimation. Where many estimates are available from previous studies, including measures of precision, researchers should use meta-estimates as these can mitigate some of the bias stemming from methodological differences in individual studies. Where new estimates or single study estimates are used in simulation models, sensitivity of the findings to different values of the elasticites should be tested, particularly for cross-price elasticities.

**Table 1. Description of data**

| | **OPEs** (n=2,749) | | **CPEs** (n=5,191) | |
|---|---|---|---|---|
| **Variables** | Obs | % | Obs | % |
| ***Study peer reviewed?*** | | | | |
| No | 2,196 | 79.9 | 3,629 | 69.9 |
| Yes | 553 | 20.1 | 1,562 | 30.1 |
| ***Country Income level*** | | | | |
| Low | 1,148 | 41.8 | 1019 | 19.6 |
| Middle | 733 | 26.7 | 948 | 18.3 |
| High | 868 | 31.6 | 3,224 | 62.1 |
| ***Region*** | | | | |
| Africa | 598 | 21.8 | 388 | 7.5 |
| Asia | 723 | 26.3 | 653 | 12.6 |
| Australasia | 58 | 2.1 | 161 | 3.1 |
| Europe | 850 | 30.9 | 1,560 | 30.1 |
| North America | 302 | 11.0 | 1873 | 36.1 |
| South America | 218 | 7.9 | 556 | 10.7 |
| ***Data type*** | | | | |
| Aggregate | 2,002 | 72.8 | 185 | 3.56 |
| Household survey data | 569 | 20.7 | 4,181 | 80.5 |
| Longitudinal survey data[a] | 178 | 6.5 | 825 | 15.89 |
| ***Data time dimension frequency*** | | | | |
| Monthly or more frequent | 306 | 11.1 | 2280 | 43.9 |
| Quarterly | 58 | 2.1 | 338 | 6.5 |
| Annual | 2,385 | 86.8 | 2,573 | 49.57 |

| | | | | |
|---|---|---|---|---|
| ***Demand system*** | | | | |
| Complete | 1,986 | 72.2 | 2181 | 42.02 |
| Conditional on food group expenditure | 383 | 13.9 | 2,098 | 40.02 |
| Conditional on food sub-group expenditure | 380 | 13.8 | 912 | 17.57 |
| ***Function type*** | | | | |
| AIDS | 738 | 26.9 | 4191 | 80.7 |
| Non AIDS | 2,011 | 73.2 | 1000 | 19.3 |
| ***Estimation type*** | | | | |
| SUR | 372 | 13.5 | 2,088 | 40.2 |
| Least Squares | 117 | 4.3 | 1,950 | 37.6 |
| Maximum Likelihood | 1,881 | 68.4 | n/a | n/a |
| Other | 97 | 3.5 | 231[b] | 4.5 |
| Not reported | 282 | 10.3 | 922 | 17.8 |
| ***How censoring in consumption data is managed?*** | | | | |
| Data aggregated or missing observations replaced by average values | 135 | 4.9 | 2132 | 41 |
| Two-step procedure | 351 | 12.8 | 1,472 | 28.4 |
| Other[c] | 34 | 1.2 | 529 | 10.2 |
| Not reported | 232 | 8.4 | 911 | 17.6 |
| Not applicable (e.g. aggregate data) | 1,997 | 72.6 | 147 | 2.8 |
| ***Which prices are used?*** | | | | |
| Retail price or price index | 159 | 5.8 | 1,542 | 29.7 |
| Unit price (adjusted to bias) | 209 | 7.6 | 896 | 17.3 |

| | | | | |
|---|---|---|---|---|
| Unit price (unadjusted to bias) | 1,130 | 41.1 | 2,092 | 40.3 |
| Other | 1,115 | 40.6 | 350 | 6.7 |
| Not reported | 136 | 5.0 | 311 | 6 |
| ***Food Group (price change)*** | | | | |
| Fruit and vegetables | 469 | 17.1 | 1,109 | 21.4 |
| Meat | 467 | 17.0 | 986 | 19 |
| Fish | 373 | 13.6 | 415 | 8 |
| Dairy | 395 | 14.4 | 610 | 11.8 |
| Eggs | 17 | 0.6 | 174 | 3.4 |
| Cereals | 376 | 13.7 | 761 | 14.7 |
| Fats and oils | 305 | 11.1 | 289 | 5.6 |
| Sweets | 47 | 1.7 | 442 | 8.5 |
| Other foods | 300 | 10.9 | 405 | 7.8 |
| ***Food Group (consumption change)***[d] | | | | |
| Fruit and vegetables | n/a | n/a | 1,140 | 22 |
| Meat | n/a | n/a | 998 | 19.2 |
| Fish | n/a | n/a | 422 | 8.1 |
| Dairy | n/a | n/a | 615 | 11.9 |
| Eggs | n/a | n/a | 179 | 3.5 |
| Cereals | n/a | n/a | 767 | 14.8 |
| Fats and oils | n/a | n/a | 306 | 5.9 |
| Sweets | n/a | n/a | 464 | 8.9 |
| Other foods | n/a | n/a | 300 | 5.8 |
| ***Mean Year*** | 2000 | | 2001 | |

[a] Studies employing scanner data were assigned one of the categories based on whether any manipulations had been done to the data (e.g. aggregation across time and/or households).

[b] Includes CPEs estimated by ML of which there were too few for a separate category

[c] Mixture of unit price and retail price, self-reported prices, comparative price levels

d CPE model only

**Table 2. Meta-regression results for own-price elasticity subsample (n=2,749)**

| **Variables** | **Categories** | **Coef.** | ***p*-value** |
|---|---|---|---|
| Publication type | Peer-reviewed | -0.004 | 0.919 |
| Income level | Middle income | 0.110 | <0.001 |
| | High income | 0.273 | <0.001 |
| Region | Africa | -0.051 | <0.001 |
| | Asia | -0.015 | 0.009 |
| | Australasia | -0.002 | 0.905 |
| | North America | -0.007 | 0.452 |
| | South America | -0.009 | 0.267 |
| Data frequency | Monthly | -0.253 | <0.001 |
| | Quarterly | -0.109 | 0.037 |
| Demand system | Complete | 0.059 | 0.127 |
| | Conditional on food sub-group expenditure | -0.021 | 0.660 |
| Function type | Non-AIDS | -0.016 | 0.853 |
| Estimation type | least squares | -0.098 | 0.198 |
| | ML | -0.065 | 0.306 |
| | Other | -0.199 | 0.001 |
| | not reported | -0.041 | 0.254 |
| Cons data censoring | Data aggregated/based on average | 0.249 | <0.001 |
| | Other | 0.338 | <0.001 |
| | Not reported | 0.226 | <0.001 |
| | Not applicable | 0.320 | <0.001 |
| Price type | Retail price | 0.093 | 0.015 |

| | | | |
|---|---|---|---|
| | Unit price (adjusted to bias) | 0.041 | 0.321 |
| | Other | 0.015 | 0.745 |
| | Not reported | 0.057 | 0.222 |
| Mean year of data | | -0.014 | 0.114 |
| Constant | | 28.65 | 0.129 |
| Food groups | | Included | |
| **Random effects parameters** | | | |
| Study ID | SD(constant) | 0.316 | |
| | SD(Residual) | 0.250 | |
| LR test vs. linear regression | $\chi^2_{(0,1)}$ = | 786.0 | <0.001 |

Note: Positive coefficients indicate less sensitive demand to changes in prices and negative coefficients more sensitive demand to changes in prices. Excluded categories: grey literature, low income country, Europe, annual data, conditional on all food expenditure demand system, AIDS or its variant function, SUR estimation, two-step approach to censored data, quality unadjusted unit price data.

**Table 3. Meta-regression results for cross-price elasticity subsample (n=5,191)**

| **Variables** | **Category** | ***Coef.*** | ***p-value*** |
|---|---|---|---|
| Publication type | Peer-reviewed | 0.028 | 0.063 |
| Income level | Middle income | n/a | n/a |
| | High income | n/a | n/a |
| Region | Africa | 0.048 | 0.103 |
| | Asia | 0.100 | <0.001 |
| | Australasia | 0.084 | 0.203 |
| | North America | 0.071 | 0.013 |
| | South America | 0.047 | 0.004 |
| Data frequency | Monthly | 0.040 | 0.012 |
| | Quarterly | 0.031 | 0.612 |
| Demand system | Complete | 0.018 | 0.195 |
| | Conditional on food sub-group expenditure | -0.005 | 0.779 |
| Function type | Non-AIDS | 0.011 | 0.37 |
| Estimation type | Least squares | -0.042 | 0.017 |
| | Other (including ML) | -0.018 | 0.471 |
| | Not reported | -0.026 | 0.216 |
| Cons data censoring | Data aggregated/based on average | 0.006 | 0.742 |
| | Other | 0.010 | 0.626 |
| | Not reported | 0.005 | 0.749 |
| | Not applicable | -0.113 | <0.001 |

| | | | |
|---|---|---|---|
| Price type | Retail price | 0.023 | 0.291 |
| | Unit price (adjusted to bias) | 0.065 | <0.001 |
| | Other | -0.074 | 0.007 |
| | not described | 0.009 | 0.696 |
| Mean year of data | | 0.001 | 0.575 |
| Constant | | -0.651 | 0.893 |
| Food group (price change) | | Included | |
| Food group (consumption change) | | Included | |
| Food group (price change)*food group (consumption change) | | Included | |
| Constant | | | |
| **Random effects parameters** | | | |
| Study ID | SD(cons) | 0.048 | |
| | SD(Residual) | 0.161 | |
| LR test vs. linear regression | $\chi^2_{(0,1)}$ = | 13.3 | <0.001 |

Note: excluded categories: grey literature, low income country, annual data, conditional on all food expenditure demand system, AIDS or its variant function, SUR estimation, two-step approach to censored data, quality unadjusted unit price data.

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## Appendix 1. Details of included studies

| Authors | Year | Country | Data frequency | Data | Function type | Estimation type | Data censoring | Price type | CPE** |
|---|---|---|---|---|---|---|---|---|---|
| Abdulai, A. | 2002 | Switzerland | Monthly | HH survey | QAIDS | SUR | Not described | Unadj. unit price | y |
| Ackah, C., Appleton, S. | 2011 | Ghana | Annual | HH survey | AIDS | SUR | Not described | Other | y |
| Adam, S. A., Sinne, S.* | 2012 | Denmark | Monthly | Longitudinal | 2-step dynamic censored AIDS | SUR | Two-step method | Unadj. unit price | y |
| Adhikari, M. et al. | 2006 | USA | Quarterly | Aggregate | LA-AIDS | SUR | N/A | Not described | y |
| Agbola, F.W. | 2003 | South Africa | Annual | HH survey | LA-AIDS | SUR | Not described | Unadj. unit price | |
| Agbola, F.W. et al. | 2003 | South Africa | Annual | HH survey | LA-AIDS | SUR | Not described | Unadj. unit price | |
| Akbay, C. et al | 2007 | Turkey | Annual | HH survey | LA-AIDS | ITSUR | Two-step method | Unadj. unit price | y |
| Akinleye, S.O.,Rahij, M.A.Y. | 2007 | Nigeria | Annual | HH survey | LA-AIDS | SURE | Not described | Unadj. unit price | |
| Alboghdady, M.A., Alashry, M.K. | 2010 | Egypt | Annual | Aggregate | LA-AIDS | RSUR | N/A | Unadj. unit price | y |
| Alderman, H.,del Ninno, C. | 1999 | South Africa | Quarterly | HH survey | AIDS | Not reported | Not described | Unadj. unit price | y |
| Alfonzo, L., Peterson, H.H. | 2006 | Paraguay | Annual | HH survey | LA-AIDS | SUR | Two-step method | Qual. adj. unit price | y |
| Allais, O. et al. | 2010 | France | Monthly | Longitudinal | AIDS | ITSUR | Aggregate/average | Qual. adj. unit price | y |
| Allais, O., Nichele, V. | 2007 | France | Monthly | Longitudinal | MS-AID | BFGS | Aggregate/average | Unadj. unit price | y |
| Allen, T. et al. | 2009 | France | Monthly | Longitudinal | LA-AI-HABIT | ITSUR | Aggregate/average | Qual. adj. unit price | y |
| Al-Shuaibi, A. | 2011 | Saudi Arabia | Annual | Aggregate | AIDS | SUR | Not described | Not described | y |
| Alviola, P., Oral, C.Jr. | 2010 | USA | Annual | HH survey | Heckman two-step model | OLS | Two-step method | Unadj. unit price | |
| Angulo, A.M. et al. | 2003 | Spain | Annual | Longitudinal | GADS | FIML | Aggregate/average | Qual. adj. unit price | y |
| Angulo, A.M., Gil, J.M. | 2006 | Spain | Annual | Longitudinal | AIDS | ITIP | Other | Qual. adj. unit price | y |

| Authors | Year | Country | Data frequency | Data | Function type | Estimation type | Data censoring | Price type | CPE** |
|---|---|---|---|---|---|---|---|---|---|
| Angulo, A.M.et al. | 2002 | Spain | Quarterly | Longitudinal | Rotterdam model | SUR-GLS | Not described | Qual. adj. unit price | |
| Anwar, A. et al.* | 2012 | Pakistan | Annual | HH survey | Rotterdam | SUR | Not described | Unadj. unit price | y |
| Balcombe, K. et al. | 1999 | Bulgaria | Monthly | HH survey | AIDS | Empirical Bayesian | N/A | Unadj. unit price | y |
| Beach, R.H., Zhen, C. | 2009 | Italy | Monthly | Aggregate | SNAP AIDS | GMM | N/A | Unadj. unit price | |
| Berges, M.E., Casellas, K.S | 2002 | Argentina | Annual | HH survey | Linear Expenditure System | SUR | Two-step method | Qual. adj. unit price | |
| Bouamra Mechemache Z. et al. | 2008 | Italy | Monthly | Longitudinal | LA-AI | ML | Not described | Retail price | y |
| Boysen, O.* | 2012 | Uganda | Annual | HH survey | QUAIDS | Iterative LS | Two-step method | Qual. adj. unit price | y |
| Brosig, S. | 2000 | Hungary | Annual | HH survey | LA-AIDS | SUR | Two-step method | Qual. adj. unit price | y |
| Brown, M. G., Jauregui, C.E.* | 2011 | US | Monthly | Aggregate | Rotterdam model | SUR | Not described | Unadj. unit price | y |
| Bunte, F., Vavra, P. | 2006 | Netherlands | Monthly | Aggregate | AIDS | SUR | N/A | Retail price | |
| Cakir, M., Balagtas, J.V. | 2010 | USA | Quarterly | Aggregate | LA-AIDS | SUR | N/A | Retail price | |
| Capacci, S., Mazzocchi, M. | 2011 | UK | Annual | HH survey | QAIDS | FIML | Two-step method | Qual. adj. unit price | |
| Caracciolo, F., Cembalo, L. | 2010 | Italy | Monthly | Longitudinal | LA-AIDS | Not reported | Two-step method | Unadj. unit price | |
| Castellon, C.E.* | 2012 | Ecuador | Annual | HH survey | AIDS | ITSUR | Two-step method | Qual. adj. unit price | y |
| Castellon, CE., et al.* | 2012 | US | Annual | HH survey | LA/EASI | SUR | Two-step method | Retail price | y |
| Coelho, A.B., et al. | 2010 | Brazil | Annual | HH survey | QAIDS | ML, nonlinear SUR | Two-step method | Unadj. unit price | y |
| Coffey, B. et al.* | 2010 | US | Monthly | Longitudinal | AIDS | EM | Other | Retail price | y |
| Conte, A. | 2006 | Egypt | Annual | HH survey | AIDS | SURE | Aggregate/average | Qual. adj. unit price | y |
| Davis, C.G. et al. | 2008 | USA | Annual | HH survey | Censored translog demand system | ML,ITSUR | Two-step method | Unadj. unit price | y |

| Authors | Year | Country | Data frequency | Data | Function type | Estimation type | Data censoring | Price type | CPE** |
|---|---|---|---|---|---|---|---|---|---|
| Davis, C.G., et al | 2011 | USA | Monthly | Longitudinal | Censored AIDS | BH | Two-step method | Not described | |
| Davis, C.G., et al. | 2007 | USA | Annual | HH survey | Censored translog demand system | ML,ITSUR | Two-step method | Unadj. unit price | y |
| Davis, C.G., et al. | 2009 | USA | Monthly | Longitudinal | Censored translog demand system | ML,ITSUR | Two-step method | Not described | |
| Dey, M.M. et al. | 2008 | Multiple | Annual | HH survey | QAIDS | Not reported | Two-step method | Not described | |
| Dharmasena, S., Capps, O.J. | 2011 | USA | Monthly | Aggregate | LA/QUAIDS | Not reported | Aggregate/average | Other | y |
| Di Giusepp, S.* | 2011 | Paraguay | Annual | HH survey | LinQuad incompleted demand system | Not reported | Not described | Qual. adj. unit price | y |
| Dong, D. et al. | 2007 | Norway | Monthly | Longitudinal | LA-AIDS | ML | Other | Qual. adj. unit price | y |
| Ecker, O., Qaim, M. | 2011 | Malawi | Annual | HH survey | QAIDS | Not reported | Two-step method | Qual. adj. unit price | |
| Elsner, K. | 1999 | Russia | Annual | HH survey | LA-AIDS | Non-linear LS | Two-step method | Qual. adj. unit price | |
| Erjavec, E., et al. | 1998 | Slovenia | Annual | HH survey | LA-AIDS | SURE | Not described | Qual. adj. unit price | |
| Fabiosa, J.F. | 2006 | Indonesia | Annual | HH survey | Double-hurdle | Likelihood fn | Two-step method | Retail price | |
| Fabiosa, J.F., Jensen, H.H. | 2002 | Indonesia | Annual | HH survey | LA-AIDS | Not reported | Two-step method | Not described | y |
| Fabiosa, J.F., Jensen, H.H. | 2003 | Indonesia | Annual | HH survey | LinQuad incomplete demand system | Not reported | Other | Not described | y |
| Fousekis, P., Revell, B.J. | 2004 | UK | Monthly | Longitudinal | Nonlinear AIDS | Not reported | N/A | Unadj. unit price | |
| Frohberg, K., Winter, E. | 2001 | Lithuania | Annual | HH survey | NQ-QES | Not reported | Not described | Not described | y |
| Garcia Y.T., et al. | 2005 | Philippines | Annual | HH survey | QAIDS | Not reported | Two-step method | Retail price | |
| Gibson, J., Rozele, S. | 2002 | Papua New Guinea | Annual | HH survey | Share-log (Deaton) | Not reported | Not described | Retail price | |
| Golan, A., et al. | 2001 | Mexico | Annual | HH survey | AIDS | GME | Other | Other | |
| Gould, B.W. | 1996 | USA | Monthly | Longitudinal | Censored demand model | ML | Two-step method | Unadj. unit price | |

| Authors | Year | Country | Data frequency | Data | Function type | Estimation type | Data censoring | Price type | CPE** |
|---|---|---|---|---|---|---|---|---|---|
| Griffith, R. et al.* | 2012 | UK | Monthly | Longitudinal | QUAIDS | GMM | Aggregate/average | Unadj. unit price | y |
| Guadelupe, B-R.J. et al. | 2010 | Mexico | Monthly | Aggregate | Simultaneous equation system | 2SLS | N/A | Retail price | |
| Gulseven, O., Wohlegant, M. | 2010 | USA | Monthly | Aggregate | Rotterdam model | Not reported | N/A | Other | |
| Gustavsen, G.W., Rickertsen, K. | 2003 | Norway | Quarterly | Aggregate | AIDS | Not reported | N/A | Retail price | y |
| Härkänen, T. et al.* | 2011 | Finland | Annual | HH survey | QAIDS | 3SLS | Not described | Retail price | y |
| Hassan, A.R. | 2012 | Columbia | Quarterly | Aggregate | Error Correction Linear AIDS | Not reported | N/A | Other | |
| Hoang, L.V. | 2009 | Vietnam | Annual | HH survey | LA-AIDS | Not reported | N/A | Qual. adj. unit price | y |
| Hoderlain, S., Mihaleva, S. | 2008 | UK | Annual | HH survey | AIDS | 3SLS/GMM | Not described | Other | |
| Hossain, F., et al. | 2001 | Latvia | Monthly | HH survey | AIDS | SUR | Not described | Retail price | y |
| Hossain, F., Jensen, H.H. | 2000 | Lithuania | Monthly | Longitudinal | LA-AIDS | OLS | Aggregate/average | Qual. adj. unit price | y |
| Huang, S-J., Show, C.-R. | 2010 | Taiwan | Monthly | Aggregate | AIDS | iterative 3SLS | N/A | Other | y |
| Huq, A.S.M.A., et al. | 2004 | Bangladesh | Annual | HH survey | LA-AIDS | Not reported | Not described | Unadj. unit price | y |
| Hutasuhut, M. et al. | 2001 | Indonesia | Annual | HH survey | LA-AIDS | Not reported | Two-step method | Not described | |
| Ishdorj, A., Jensen, H.H. | 2008 | USA | Monthly | Longitudinal | Censored AIDS | Bayesian | Other | Unadj. unit price | |
| Islam, M.R. et al. | 2007 | Bangladesh | Annual | HH survey | LA-AIDS | OLS | Not described | Unadj. unit price | y |
| Ismail, S.Z., Lofti, G.R. | 2007 | Egypt | Annual | Aggregate | Barton mixed model/AIDS | Not reported | Not described | Not described | y |
| Jabarin, A.S., Al-Karablieh, E.K. | 2011 | Jordan | Annual | HH survey | LA-AIDS | ITSUR | Two-step method | Not described | |
| Jaffry, S., Brown, J. | 2008 | UK | Monthly | Aggregate | Dynamic AIDS | Not reported | N/A | Unadj. unit price | |
| Klonaris, S., Karagiannis, G. | 2002 | Greece | Annual | HH survey | LA-AIDS | Not reported | Two-step method | Qual. adj. unit price | |

| Authors | Year | Country | Data frequency | Data | Function type | Estimation type | Data censoring | Price type | CPE** |
|---|---|---|---|---|---|---|---|---|---|
| Kuchler, F. et al. | 2010 | USA | Monthly | Aggregate | LA-AIDS | ISUR | N/A | Unadj. unit price | |
| Kumar, P., Dey, M.M. | 2004 | Canada/ India | Annual | HH survey | QAIDS | Not reported | Aggregate/average | Other | |
| Lazaridis, P. | 2003 | Greece | Annual | HH survey | LA-AIDS | SURE | Two-step method | Qual. adj. unit price | |
| Le, C.Q. | 2008 | Vietnam | Annual | HH survey | AIDS | OLS | Not described | Retail price | |
| Lecocq, S., Robin, J.-M. | 2006 | France | Quarterly | Longitudinal | QAIDS | ITSUR | Aggregate/average | Unadj. unit price | |
| Leffler, K.K. et al.* | 2012 | US | Annual | HH survey | EASI | SUR | Aggregate/average | Qual. adj. unit price | y |
| Lema, D., et al. | 2007 | Paraguay/ Bolivia | Annual | HH survey | LinQuad | ISUR | Two-step method | Unadj. unit price | y |
| Lin, B.H., et al. | 2008 | USA | Annual | HH survey | Translog model | Not reported | Two-step method | Unadj. unit price | |
| Lin, B.H., et al. | 2011 | USA | Monthly | Aggregate | AIDS | ITSUR | N/A | Unadj. unit price | y |
| Llanto, G.M. | 1996 | Philippines | Annual | HH survey | QAIDS | ITSSUR | Not described | Not described | y |
| Lopez, J.A, Malaga, J. E. | 2009 | Mexico | Annual | HH survey | Two-step censored demand model | ML | Two-step method | Unadj. unit price | y |
| Luchini, S, R., et al. | 2001 | Bulgaria | Monthly | Aggregate | AIDS | SUR | N/A | Not described | y |
| Ma, H. et al. | 2003 | China | Annual | Aggregate | LA-AIDS | ITSUR | N/A | Other | |
| Maynard, L.J. | 2000 | USA | Monthly | Aggregate | LA-AIDS | ITSUR | N/A | Retail price | |
| Maynard, L.J., Liu, D. | 1999 | USA | Monthly | Aggregate | LA-AIDS | Not reported | N/A | Unadj. unit price | |
| Mazzocchi, M. | 2004 | Italy | Monthly | Aggregate | AIDS | SUR | Aggregate/average | Retail price | y |
| Meyerhoefer, C.D., et al. | 2005 | Romania | Monthly | Longitudinal | Continuous/censored commodity demand system | GMM | Other | Qual. adj. unit price | y |
| Minot, N., Goletti, F. | 2000 | Vietnam | Annual | HH survey | LA-AIDS | Not reported | Not described | Other | |
| Monnet Benoit, P.G., Souza-Posa, A. | 2011 | Cote d'Ivoire | Annual | HH survey | LA-AIDS | 2SLS | Aggregate/average | Qual. adj. unit price | y |
| Moschini, G., Rizzi, P.L. | 2007 | Italy | Monthly | Aggregate | NQ Mixed Demand System | ML | N/A | Qual. adj. unit price | |

| Authors | Year | Country | Data frequency | Data | Function type | Estimation type | Data censoring | Price type | CPE** |
|---|---|---|---|---|---|---|---|---|---|
| Moschini, G., Rizzi, PL. | 2005 | Italy | Monthly | Aggregate | Stone-Geary Mixed Demand Model | ML | N/A | Qual. adj. unit price | |
| Mudassar, K. et al.* | 2012 | Pakistan | Annual | HH survey | LA/AIDS | ITSUR | Not described | Unadj. unit price | y |
| Muhammad, A., et al. | 2011 | Multiple | Annual | Aggregate | Florida-Slutsky | ML | N/A | Other | |
| Mutondo, J.E, Henneberry, S.R. | 2007 | USA | Quarterly | Aggregate | Rotterdam model | ITSUR | N/A | Unadj. unit price | |
| Niimi, Y. | 2005 | Vietnam | Annual | HH survey | LA-AIDS | SUR | Not described | Retail price | |
| Okrent, A.M., Alston, J.M. | 2011 | USA | Monthly | Aggregate | FD-LAIDS | ITSUR | Aggregate/average | Retail price | y |
| Okrent, A.M., Alston, J.M.* | 2012 | USA | Monthly | HH Survey | GODDS | GLS | Aggregate/average | Retail price | y |
| Ozer, H. | 2003 | Turkey | Annual | HH survey | Linear Expenditure System | SUR | Not described | Retail price | y |
| Peterson, H.H., Chen Y. | 2005 | Japan | Monthly | Aggregate | Rotterdam model | Not reported | N/A | Retail price | y |
| Piggot, N.E., et al. | 2007 | USA | Monthly | Aggregate | Generalized AIDS | ITSUR | N/A | Unadj. unit price | |
| Pintos-Payeras, J.A. | 2009 | Brazil | Annual | HH survey | AIDS | Not reported | Not described | Retail price | y |
| Pittman, G.F. | 2004 | USA | Annual | HH survey | LA-AIDS | SUR | Two-step method | Unadj. unit price | y |
| Pofahl, G.M., et al. | 2005 | USA | Annual | HH survey | QAIDS | ITSUR | N/A | Unadj. unit price | y |
| Pomboza, R. Mbaga, M. | 2007 | Canada | Annual | HH survey | AIDS | SUR | Aggregate/average | Unadj. unit price | y |
| Pruitt, J.R., Raper, K.C. | 2010 | USA | Monthly | Aggregate | AIDS | GMM | Not described | Retail price | |
| Quagrainie, K. | 2003 | USA | Monthly | Aggregate | Dynamic AIDS | Non-linear procedure in SHAZAM | N/A | Other | |
| Radwan, A. et al. | 2009 | Spain | Monthly | Aggregate | Generalized AIDS | Not reported | N/A | Retail price | |
| Radwan, A., et al. | 2008 | Spain | Monthly | Aggregate | Generalized AIDS | Largest likelihood function value | N/A | Retail price | |
| Ragab, M.A.S., et al. | 2008 | Egypt | Annual | Aggregate | LA-AIDS | 3SLS | Not described | Not described | y |

| Authors | Year | Country | Data frequency | Data | Function type | Estimation type | Data censoring | Price type | CPE** |
|---|---|---|---|---|---|---|---|---|---|
| Ramadan, R., Thomas, A. | 2011 | Egypt | Annual | Aggregate | Mixed demand model | Non-linear SUR | Aggregate/average | Other | y |
| Raper, K.C. | 2002 | USA | Annual | HH survey | Linear Expenditure System | Non-linear SUR | Two-step method | Retail price | |
| Razzaque, A. et al. | 1997 | Bangladesh | Annual | HH survey | Food Characteristics Demand System (FCDS) | Not reported | Not described | Other | y |
| Regorsek, D., Erjavec, E. | 2007 | Slovenia | Annual | HH survey | LA-AIDS | System linear regression | Aggregate/average | Unadj. unit price | |
| Revoredo-Giha, C., et al. | 2009 | Scotland | Monthly | Aggregate | LA-AIDS | SURE | N/A | Unadj. unit price | |
| Rickertsen, K., Kristofersson, D. | 2003 | Norway | Annual | Aggregate | LA-AIDS | 3SLS | N/A | Other | |
| Rickertsen,K. | 1998 | Norway | Annual | Aggregate | AIDS | SUR | N/A | Other | |
| Santarossa, J.M, Mainland, D.D. | 2003 | UK | Monthly | Longitudinal | AIDS | Not reported | Aggregate/average | Unadj. unit price | y |
| Schmit, T.M, et al. | 2002 | USA | Monthly | Longitudinal | Two-step censored demand model | ML | Two-step method | Unadj. unit price | |
| Sckokai, P, et al. | 2009 | Italy | Monthly | Aggregate | AIDS | GMM | Aggregate/average | Unadj. unit price | |
| Seale, J., et al. | 2003 | Multiple | Annual | Aggregate | Florida-Slutsky | ML | N/A | Unadj. unit price | |
| Shirota, R., Sonoda, D.Y.* | 2012 | Brazil | Annual | HH survey | AIDS | GLS | Not described | Not described | y |
| Smed, S., et al. | 2007 | Denmark | Monthly | Aggregate | AIDS | ML | Aggregate/average | Unadj. unit price | |
| Souza, G.S., et al. | 2008 | Brazil | Annual | Aggregate | Partial Equilibrium Model | 3SLS | N/A | Other | |
| Stockton, C., Capps, O. | 2005 | USA | Annual | HH survey | Censored AIDS | OLS | Two-step method | Unadj. unit price | y |
| Taniguchi, K., Chern, W.S | 2000 | Japan | Monthly | HH survey | AIDS | ITSUR | Two-step method | Other | y |
| Tekguc, H.* | 2011 | Turkey | Annual | HH survey | LA/AIDS | FGLS | Two-step method | Not described | y |
| Tey, S.Y., et al. | 2008 | Malaysia | Annual | HH survey | LA-AIDS | Not reported | Not described | Not described | |
| Tey, S.Y., et al. | 2008 | Malaysia | Annual | HH survey | LA-AIDS | ML | Two-step method | Not described | |

| Authors | Year | Country | Data frequency | Data | Function type | Estimation type | Data censoring | Price type | CPE** |
|---|---|---|---|---|---|---|---|---|---|
| Thiele, S. | 2008 | Germany | Annual | HH survey | LA-AIDS | SUR | Two-step method | Qual. adj. unit price | |
| Thiele, S. | 2010 | Germany | Annual | HH survey | LA-AIDS | SUR | Two-step method | Qual. adj. unit price | y |
| Tiffin, R. et al.* | 2011 | UK | Annual | HH survey | AIDS | SUR | Other | Unadj. unit price | y |
| Tiffin, R., Arnoult, M. | 2010 | UK | Annual | HH survey | AIDS | Bayesian | Other | Unadj. unit price | y |
| Tinooco, J.R., et al. | 2011 | Mexico | Monthly | Aggregate | AIDS | SUR (SYSLIN/SUR) | N/A | Retail price | y |
| Turk, J., Erjavec, E. | 2001 | Slovenia | Annual | HH survey | LA-AIDS | SURE | Not described | Unadj. unit price | y |
| ul Haq, Z. et al. | 2008 | Pakistan | Annual | HH survey | LA-AIDS | ITSUR | Not described | Not described | |
| Ulimwengu, J.M. et al. | 2009 | Ethiopia | Annual | HH survey | AIDS | SUR | Not described | Not described | y |
| Ulimwengu, J.M., Ramadan, R. | 2009 | Uganda | Annual | HH survey | AIDS | Not reported | Not described | Not described | |
| Ulubasoglu, M. et al. | 2010 | Australia | Quarterly | HH survey | LA-AIDS | Not reported | Two-step method | Retail price | y |
| Verbeke, W., Ward, R. W. | 2001 | Belgium | Monthly | Longitudinal | AIDS | Not reported | N/A | Not described | |
| Weliwita, A., et al. | 2003 | Tanzania | Annual | HH survey | LA-AIDS | Nonlinear ITSUR | Two-step method | Unadj. unit price | y |
| Yeboah, G., Maynard, L.J. | 2004 | Japan | Monthly | Aggregate | Rotterdam model | SUR | N/A | Retail price | y |

*only cross-price elasticities are extracted from these studies as search for publications estimating cross-price elasticities was done separately and with a later end date. While own-price elasticity estimates are available from these studies, these are not included to avoid bias as this would exclude studies not presenting cross-price elasticities and dating beyond August 2011.

**Studies present cross-price elasticities

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