6-1996

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Abstract
How can economists estimate consumption parameters by income classes? The approach used in this paper combines price elasticities estimated from aggregated market data and income-class-specific elasticities derived from household expenditure surveys using Slutsky relationships to calculate income-class-specific price elasticities. The approach was applied to estimate income-class-specific price elasticities for major agricultural commodities consumed in Jamaica.

Disciplines
Agricultural and Resource Economics | Agricultural Economics | Behavioral Economics | Public Policy

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Working Paper 96-WP 159
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Funding for this research was provided through the U.S. Agency for International Development under contract number FAO-0800-C-00-4070-00.

ABSTRACT

This paper develops an approach to estimate consumption parameters by income classes. It combines price elasticities estimated from aggregated market data and income-class-specific income elasticities derived from household expenditure surveys using Slutsky relationships to calculate income-class-specific price elasticities. The approach was applied to estimate income-class-specific price elasticities for major agricultural commodities consumed in Jamaica. The importance of these elasticities for food policy analysis is demonstrated by using an econometric simulation model framework. Food policy analysis based on aggregate consumption parameters may introduce bias into the results. This is particularly true in the case of developing countrieswhere disparity between low- and high-income groups is very high.

USING INCOME CLASSES TO ESTIMATE CONSUMPTION PARAMETERS FOR FOOD POLICY ANALYSIS

Food policy analysis that links nutrition objectives to macroeconomic policies and performance requires a matrix of price and income elasticities that is income-strata-specific (Timber and Alderman, 1979). But from a practical point of view, food policy analysis is conducted using price and income elasticities for aggregated income classes, estimated using complex theory and sophisticated econometrics techniques. Use of these parameters based on per capita or per household terms induces aggregation bias into the result. The opportunity for consumers at different levels to behave differently with respect to economic parameter variations (e.g.,price changes) is obliterated in the data aggregation process.

In order to avoid aggregation bias, cross-section analysis has been used to understand more about how household decision parameters vary at different income levels. Income-class-specific income elasticities are easy to derive from cross-section consumption surveys, whereas the use of cross-section data to derive income-class-specific price elasticities has been extremely limited for fairly obvious reasons: most of the household surveys are collected at a particular time when the prices they face are given (Timmer and Alderman, 1979). One way to avoid this problem is either to have a cross-section panel of consumers whose consumption expenditures are recorded over time, or to draw a large sample with enough geographical and temporal diversity to capture significant variation in relevant variables.

Studies such as Timmer and Alderman (1979), Alderman (1986), Fan et al.(1994), and Fan et al. (1995) have used cross-section panel data to estimate income-specific price elasticities for various policy analysis purposes. But in reality it is very hard to estimate because data series are short and not too accurate, as is data typically received from many developing countries. Of more importance, many developing countries may not have cross-section panel data collected over time or a sample large enough to cover sufficient geographical and temporal diversity to represent the entire population. Instead, they may have cross-section household expenditure surveys collected at a specific time. In those cases, estimation of income-specific price elasticities is out of question.

This paper uses an approach that combines price elasticities estimated from aggregated market data and income-class-specific income elasticities derived from household expenditure surveys using Slutsky relationships to calculate income-class-specific price elasticities. This approach may be extended to estimate price elasticities pertaining to any demographic characteristics such as location, gender, or profession, as long as group-specific income elasticities can be estimated.

Derivation of Aggregate and Income-Specific Price Elasticities

In this study, we used aggregate market data and household expenditure survey information for Jamaica to estimate income-class-specific price elasticities for various commodities. Since there is a lag in consumer response to price changes, a dynamic version of structural Linear Approximation Almost Ideal Demand System (LA/AIDS) developed by Wickens and Bruesch was used in this study to estimate price elasticities from aggregated market data. The dynamic structural model, developed by Wickens and Bruesch (1989), may be specified as:

$$\sum_{j=0}^{m} W_{t-j} A_j = \sum_{j=0}^{n} X_{t-j} B_j + E_t \tag{1}$$

where W = $(w_1, w_2, ..., w_s)$ is a vector of s budget shares and the jth commodity share is $w_j = (p_j\, q_j / m)$, q_j is the quantity demanded of the jth commodity, m is the group expenditure on m commodities. X = (1, $Ln(p_1)$, $Ln(p_2)$, ..., $Ln(p_s)$, Ln(m/P); p_j is the nominal price of the jth commodity; and Ln(P) is the Stone Price Index defined as $Ln(P) = \sum w_j Ln(p_j)$. E_t is a vector of stochastic error terms distributed as i.i.d. $(0,\Omega)$. B_j is an $(s\times(s+2))$ coefficient matrix. A_j = (1, 2, ..., m) is an s × s coefficient matrix assumed to be diagonal to avoid perfect collinearity among regressors arising from the singular equation system ($\sum w_{j,\, t-i} = 1\ \forall$ i). As a result of the diagonal A_j, each budget share depends only on its lags and not on the lags of other budget shares.

The structural system (1) is transformed into the reduced form

$$W_t = \sum_{j=1}^{m} W_{t-j} C_j + \sum_{j=0}^{n} X_{t-j} D_j + U_t \tag{2}$$

where the transformed coefficient matrixes are $C_j = -A_j A_0^{-1}$, $D_j = B_j A_0^{-1}$, and $U_t = E_t A_0^{-1}$.

The reduced form equation is adequate to obtain the estimates of C_j and $D_j\ \forall$ j, and further computations are required to derive the long-run parameter matrix Φ using the formula,

$$\Phi = \frac{\sum_{j=0}^{n} D_j}{I - \sum_{J=0}^{m} C_j} . \tag{3}$$

However, computing the long-run parameters from the short-run estimates will not allow easy imposition of the theoretical restrictions such as homogeneity, symmetry, and adding up on the long-run parameters. Thus, it is imperative to estimate the long-run parameters directly so that theoretical restrictions can be imposed on these parameters. This can be done by transforming and reparameterizing the reduced form (2) into an observationally equivalent formulation to allow for direct estimation of the long-run parameter (Φ). This reparameterized formulation is written as:

$$W_t = \sum_{j=1}^{m} \Delta_j W_t F_j + X_t \Phi + \sum_{j=0}^{n} \Delta_j X_t G_j + V_t , \tag{4}$$

where $F_j = C_j H$, $G_j = D_j H$, $V_j = U_j H$, and the $\Delta_j W_t = W_t - W_{t-j}$.

A single equation of (4) representing the i^{th} budget share in the reparameterized formulation can be written as

$$\begin{aligned} w_{it} = & \sum_{j=0}^{m} f_{iij} \Delta_j w_{it} + \phi_{i0j} + \sum_{k=1}^{s} \sum_{j=1}^{n} \phi_{ikj} Ln(p_{k,t-j}) + \sum_{j=0}^{n} \phi_{i(s+1)j} Ln(\frac{M_t}{P_t}) \\ & \sum_{k=1}^{s} \sum_{j=1}^{n} g_{ikj} \Delta_j Ln(p_{k,t-j}) + \sum_{j=1}^{n} g_{i(s+1)j} \Delta_j Ln(\frac{M_{t-j}}{P_{t-j}}) + v_{it} , \end{aligned} \tag{5}$$

where f_{iij} is the element in the i^{th} row and i^{th} column of F_j coefficient matrix. ϕ_{ik} is the k^{th} element in the i^{th} column of ϕ, and g_{kij} is the k^{th} element in the i^{th} column of G_j.

Anderson and Blundell (1983) argued that, since demand theory is derived under equilibrium conditions, it is more likely that theoretical restrictions would hold in the long run rather than the short run. The model in (4) allows us to impose the demand restrictions implied by the axioms of preference in demand theory (i.e., adding up, homogeneity, and symmetry) only on the long-run parameters. The demand restrictions in this formulation (equivalent to those originally derived by Deaton and Muellbauer, 1980) are

Adding up,

$$\sum_{i=1}^{s} \phi_{i0} = 1, \quad \sum_{i=1}^{s} \phi_{ik} = 0, \quad \forall\ K=1,\ \ldots,\ S+1\ ; \tag{6}$$

Homogeneity,

$$\sum_{i=1}^{s} \phi_{ik} = 0\ ; \tag{7}$$

and

Symmetry

$$\phi_{jk} = \phi_{kj}\ \forall\ k,j = 1, \ldots, s\ , \tag{8}$$

where ϕ_{i0} is the intercept and ϕ_{ik} (k= 1, ..., s) are parameters of prices.

Equation (4) and this set of restrictions can be used to directly estimate the long-run parameter. However, OLS estimation would yield biased and inconsistent parameter estimations because in equation (4), $\Delta_j W_t$ is correlated to V_t. As shown by Philips and Wickens (1978), simultaneity problems can be corrected by maximum likelihood estimation of system (4).

Conditional Elasticities from Time Series Data

The group or conditional (long-run) expenditure and price elasticities are given by Green and Alston (1991) and Foster, Green, and Alston (1990).

$$\eta^{G}_{iM} = \frac{\phi_{iM}}{W_i} + 1 \tag{9}$$

$$\eta^{G}_{ij} = (\phi_{ij} - \phi_{iM} W_j)/W_i - \delta_{ij} \tag{10}$$

where δ = the kronecker delta, subscript M = the expenditure, superscript G = group or conditional, and W = the predicted budget share.

As suggested by Kesavan et al. (1993), the total or unconditional demand elasticities are preferred to conditional elasticities for policy analysis. Furthermore, conditional elasticities are difficult to compare with other studies. The procedure includes the estimation of group expenditure (M).

$$\log(M_t) = \lambda_0 + \Gamma_1 \log GP_{1t} + \Gamma_2 \log GP_{2t} + \Gamma_3 \log POF_t + \Gamma_4 \log PNF_t \\ + \Gamma_y \log PCE_t + \Gamma_T \, Trend + h_t \tag{11}$$

where

M = the group expenditure (in per-capita terms),

GP_1 and GP_2= group price indexes (e.g. meat group),

POF = the price of foodless group,

PNF = price of nonfood goods,

PCE = the per-capita personal consumption expenditure,

Trend= a time trend variable.

The group price indexes are derived as

$$\log(GP_{1t}) = \sum_{i=1}^{s} w_{it} \log(P_{it}) \ , \tag{12}$$

and

$$\log(GP_{2t}) = \sum_{i=1}^{s} W_{it} \, \eta^{G}_{it} \log(P_{it}) \ , \tag{13}$$

where W_i = group budget share of the i^{th} commodity, and η^{G}_{it} is the group expenditure elasticities.

The first index, GP_1, is the geometric weighted index of prices, where the weights are based on within-group budget shares, whereas the second index, GP_2, is weighted by within-group expenditure elasticities. Thus, GP_1 reflects the "substitution" effects of a within-group price change, while GP_2 reflects the "expenditure" effects from changes in relative prices within the group.

As in Kesavan et al. (1993), the total or unconditional price and income elasticities can be calculated from the long-run conditional or group elasticities of the Generalized Dynamic Almost Ideal Demand System (GD/AIDS) and the estimates of first-stage allocation (equation 6) using the formulas

$$\eta^{T}_{iM} = \eta^{G}_{iM} E_y \ , \tag{14}$$

and

$$\eta^{T}_{ij} = \eta^{G}_{ij} + \eta^{G}_{iM} (E_1 W_j + E_2 W_j \eta^{G}_{jM}) \tag{15}$$

The unconditional elasticities derived from the time series data provide an aggregate measure of the responsiveness of consumers. These estimates can be enriched with more desegregated information from household expenditure surveys that can provide measures of differential responsiveness based on income, location, and other household characteristics. What follows is a proposed procedure for constructing new elasticity estimates by merging information from time series and from the household expenditure surveys. The starting equation is the Slutsky decomposition of elasticity into the substitution and income effects. That is, the elasticity from the time series can be decomposed into

$$\eta^{u}_{ij} = \eta^{*}_{ij} - w_j \ \eta^{u}_{i} \tag{16}$$

where η^{*}_{ij} is the compensated or Hicksian price elasticity and w_j is the proportion of total expenditure spent on commodity j.

The key assumption in this methodology is that differential responsiveness of consumers is attributed completely to the income effect. From these Slutsky relationships, a Hicksian elasticity, which is assumed to be constant across households, can be estimated. That is,

$$\eta_{ij}^{*} = \eta_{ij}^{u}+w_j\ \eta_i^{u} = \eta_{ijh}^{*}\ \forall h \tag{17}$$

where h is the household index (for income class, quartiles one to four). With the elasticities from the time series data, household-specific income elasticities and expenditure share by commodity from the household expenditure survey, a set of elasticity estimates by household category can be constructed using the following formula

$$\eta_{ijh} = \eta_{ijh}^{*} - w_{jh}\ e_{ih}\ . \tag{18}$$

The additional information provided by the household expenditure survey is the differential income elasticity for each commodity across households (e_{ih}) and the share of each commodity across households (w_{jh}).

Empirical Example

This method was applied to Jamaican time series and household expenditure survey data. The time series data covers 1972 to 1993. Two broad group of commodities, meat and crops, were included for analysis. The meat group included beef, pork, and chicken, and the crop group was wheat, rice, sugar, and soyoil. Time series data such as consumption and prices of these commodities, population, various price indexes (such as food and nonfood) were collected from various sources. Consumption series data were approximated by the disappearance series, which is derived as a residual in an accounting identity of the sources and uses of the food. Sources of food include current production, imports, and beginning inventory. The uses of food (excluding human consumption) are feed, exports, and ending inventory. Human consumption is calculated by deducting nonfood uses from the source of supply. This approach was used for both meat and crops.

Both the sources and uses data are collected from *Economic and Social Survey of Jamaica* (various issues) published by the Planning Institute of Jamaica. The retail prices are collected from *Consumer Price Index* published by the Statistical Institute of Jamaica. Most of the macro variables are collected from *The Statistical Yearbook of Jamaica* by the Statistical Institute of Jamaica.

Data from household expenditure surveys were needed to calculate the income elasticities for different income groups and the share of consumption for each commodity for different income groups. This information was obtained from a household expenditure survey conducted in 1984.

Estimation and Results

Both meat and crop demand were specified and estimated using dynamic AIDS as described in the previous section. Lag values of the expenditure shares, price,s and trends were included to capture dynamic adjustment of consumers. The theoretical demand properties were imposed only on the long-run parameters and estimated as a system of equations using Iterative Three-Stage Least Squares. This method gives maximum likelihood estimates at the point of convergence. Estimation was accomplished through SAS and RATS.

Table 1 shows the estimates of the meat demand system. The adequacy of the estimated model is reflected by a number of statistics. The estimated model displays all the theoretical demand properties since these were imposed in the estimation. Many of the long-run parameters have coefficient estimates that are significant. Also lagged regressors and trends are significant, suggesting dynamic adjustment of consumers. Long-run parameters are used to calculate conditional expenditure and price elasticities.

In the next step, we estimated group expenditure (equation 6) using Ordinary Least Squares. The estimated model is reasonable in explanatory power and magnitude of the coefficients.[1] Based on the long-run conditional elasticities and the coefficients of group expenditure equations, the total or unconditional price and expenditure elasticities are calculated using the formulas specified earlier (equations 14 and 15).

The estimated unconditional price and expenditure elasticities for meat and crop groups are reported in Tables 3 and 4. In terms of magnitude of own price response, pork tops the list followed by beef and chicken. This is logical because per capita chicken consumption is much higher than beef and pork consumption and thus, any small percentage change is very high in absolute terms. For example, in absolute terms, a 1 percent change in chicken consumption is approximately equal to a 30 percent change for pork consumption and 5 percent change in beef consumption. Positive and negative cross-price elasticities in the meat and crop groups suggest that both substitute and complement relationships exists within each group.

The unconditional price elasticities are then converted to Hicksian price elasticities using the Slutsky equation specified in equation 16. The Hicksian price elasticities for meat and crop groups are reported in Tables 5 and 6. These Hicksian price elasticities, along with information from the household expenditure

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[1]Results are available from the authors upon request.

survey, are used to calculate differential response based on income. In this study, the entire population is divided into four different income quartiles. For each income group, income elasticities and shares of each commodity are calculated from the household expenditure survey.[2] Tables 7 and 8 report the differential price response of four income groups for the meat and crop groups. As expected, the overall results suggest that the lower income group is more responsive to price change than the higher income group. A similar approach can be used to estimate price elasticities for various demographic groups such as location, profession, and family size.

Further, the importance of these disaggregated elasticities for food policy analysis is demonstrated when they are used in econometric simulation model framework. One interesting aspect of the model is that per capita consumption of commodities is translated into intake of major nutrients. The formula used to convert consumption to nutrients is:

$$TN^{h}{}_{i} = \sum_{j=1}^{n} \beta_{ij} Q^{h}{}_{a,\ldots,j} \ , \tag{19}$$

where TN^h_i is the total nutrient intake of the i^{th} nutrient for the h^{th} household category with $h = Q_{1\text{-}4}$. β_{ij} is the proportion of the i^{th} nutrient in the j^{th} commodity, as consumed. The vector of the micro- and macronutrients includes energy, protein, fat, carbohydrate, iron, calcium, vitamin, thiamine, and riboflavin.

Total intake is then compared with recommended daily allowances (RDA) to determine shortfalls or excesses from policy changes. The ratio of total intake of nutrient i to its corresponding RDA is expressed by

$$ADQ^{h}{}_{i} = \frac{TN_i^{h}}{RDA_i} * 100 \ . \tag{20}$$

As this ratio approaches 100 percent, nutritional status improves toward the optimum intake level for the household Q_i. Further description of the model's structure can be obtained from the authors upon request.

Using the simulation model, we analyzed the impact of reducing import tariffs of selected commodities on consumption and nutritional intakes of different income groups for 1995 to 2005. Figures 1

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[2]More information on these calculations is available upon request.

and 2 report the differential response in consumption of two major commodities for the highest (quartile 4) and lowest (quartile 1) income groups due to import tariffs reduction. Figure 1 shows the percentage change in per capita chicken consumption for quartile one and four households. On average, chicken consumption increased by 16 percent for quartile 1 compared to 8 percent for quartile 4. Similarly, Figure 2 presents the percentage change in per capita wheat consumption for both these quartiles. Similar to chicken consumption results, lower income households have a higher percentage increase in wheat consumption when compared with the higher income groups (on average, chicken consumption for quartile 1 increased by 4 to 5 percent compared with less than 1 percent for quartile 4).

Figures 3 and 4 report percentages of recommended of two major nutrients such as energy and protein, with both import tariffs and reduced import tariffs. By reducing import tariffs, energy intake of quartile 1 increased from 70 to 78 percent of the recommended level over years, whereas for quartile 4, it increased from 77 to 79 percent over the same time. Similarly, Figure 4 shows the percentage of recommended protein intake for quartiles 1 and 4, with both import tariffs and reduced import tariffs. Protein intakes increased from about 90 to 100 percent of the recommended level for quartile 1. For quartile 4, protein intakes, which was little above recommended level before import tariffs reduction increased little more by reducing import tariffs.

Overall, the results suggest that lower income groups rather than higher income households benefit most from reduced import tariffs. Of more inportance, disparity in consumption and nutritional status between the high and low income groups reduced significantly with these policy changes. This information is vital for policymakers in reformulating policies to reach the targeted groups.

Conclusion

This paper estimates consumption parameters for different income classes using a method that combines both time series and household expenditure survey data. This approach is more useful in the case where panel data are not available to estimate consumption parameters for income classes directly, as in most developing countries. Further, this approach was used with Jamaican time series and household expenditure survey data to estimate demand elasticities for different income groups. These demand parameters were then used in a simulation model framework to analyze the impact of variable import tariff reduction on the consumption and nutritional intake of different income groups. The results suggest that reducing selected import tariffs improves the nutritional status of lower income households much more than that it does those of higher income households. This type of information enables policymakers to direct policies toward the targeted groups of the society.

Table 1. Maximum Likelihood Estimates of Meat Demand

Variable	Coefficient	Standard Error
Dependent		
Share of Beef		
Independent		
Constant	0.834	0.172
Log of Price of Beef	-0.077	0.022
Log of Price of Chicken	-0.033	0.020
Log of Real Expenditure	-0.071	0.039
First Difference of Beef Share	0.276	0.042
Second Difference of Beef Share	0.252	0.031
First Difference of Price of Beef	0.054	0.049
First Difference of Price of Chicken	0.014	0.023
First Difference of Price of Pork	-0.050	0.042
Trend	-0.013	0.001
Dependent		
Share of Chicken		
Independent		
Constant	-0.336	0.185
Log Price of Chicken	0.093	0.022
Log of Real Expenditure	0.179	0.042
First Difference of Chicken Share	0.222	0.039
Second Difference of Chicken Share	0.165	0.027
First Difference Price of Beef	-0.071	0.053
First Difference Price of Chicken	-0.034	0.025
First Difference Price of Pork	0.073	0.044
Trend	0.012	0.001

Table 2. Parameter Estimates of Crop Demand

Variable	Coefficient	Standard Error
Dependent		
Share of Wheat Flour		
Independent		
Constant	-0.362	0.230
Log of Retail Price of Flour	0.081	0.022
Log of Retail Price of Rice	0.053	0.016
Log of Retail Price of Sugar	-0.088	0.020
Log of Retail Price of Soyoil	-0.022	0.009
Log of Real Expenditure	0.121	0.040
First Difference of Wheat	0.352	0.044
Trend	0.003	0.002
Dependent		
Share of Rice		
Independent		
Constant	0.122	0.155
Log Retail Price of Rice	0.012	0.032
Log Retail Price of Sugar	-0.062	0.017
Log Retail Price of Soyoil	0.017	0.009
Log of Real Expenditure	0.003	0.027
First Difference of Rice	0.627	0.108
Trend	0.002	0.001

Table 2. (Continued)

Variable	Coefficient	Standard Error
Dependent		
Share of Sugar		
Independent		
Constant	1.916	0.334
Log Retail Price of Sugar	0.190	0.030
Log Retail Price of Soyoil	-0.010	0.009
Log of Real Expenditure	-0.264	0.059
First Difference of Sugar	0.236	0.079
Trend	-0.006	0.002
Dependent		
Share of Soyoil		
Independent		
Constant	-0.235	0.129
Log Retail Price of Soyoil	0.011	0.008
Log Real Expenditure	0.040	0.022
First Difference of Soyoil	0.591	0.129
Trend	0.002	0.001

Table 3. Unconditional Price and Expenditure Elasticities for Meat

	Beef	Chicken	Pork	Expenditure
Beef	-0.961	0.299	0.355	0.672
Chicken	0.082	-0.502	-0.103	1.130
Pork	-1.185	0.015	-1.267	0.146

Table 4. Unconditional Price and Expenditure Elasticities for Crops

	Wheat Flour	Rice	Sugar	Soy oil	Cornmeal	Expenditure
Wheat Flour	-0.527	0.101	0.116	0.121	0.064	0.383
Rice	0.233	-0.614	-0.138	0.141	0.289	0.261
Sugar	0.336	-0.022	-0.439	0.050	0.025	0.275
Soy oil	0.380	0.291	-0.189	-0.597	-0.091	0.425
Cornmeal	-0.005	0.483	-0.340	-0.005	-0.282	0.519

Table 5. Hicksian Price Elasticities for Meat

	Beef	Chicken	Pork
Beef	-0.528	0.349	0.179
Chicken	0.267	-0.316	0.049
Pork	0.512	0.182	-0.694

Table 6. Hicksian Price Elasticities for Crops

	Wheat Flour	Rice	Sugar	Soyoil	Cornmeal
Wheat Flour	-0.480	0.136	0.160	0.130	0.079
Rice	0.266	-0.589	-0.106	0.147	0.299
Sugar	0.355	-0.008	-0.421	0.053	0.031
Soyoil	0.457	0.350	-0.115	-0.582	-0.067
Cornmeal	0.066	0.541	-0.268	-0.040	-0.260

Table 7. Differentiated Elasticities in Meat Products by Income Groups

	Beef	Poultry	Pork
Quartile 1			
Beef	-1.202	-0.566	0.234
Poultry	0.131	-0.779	-0.079
Pork	0.947	-0.651	-1.387
Quartile 2			
Beef	-1.186	-0.244	0.251
Poultry	0.177	-0.558	-0.053
Pork	0.926	-0.480	-1.391
Quartile 3			
Beef	-1.243	-0.148	0.263
Poultry	0.175	-0.496	-0.041
Pork	0.912	-0.361	-1.369
Quartile 4			
Beef	-1.323	-0.001	0.281
Poultry	0.174	-0.430	-0.028
Pork	0.868	-0.243	-1.352

Table 8. Differentiated Elasticities in Crop Products by Income Groups

	Wheat	Rice	Sugar	Soyoil	Cornmeal
Quartile 1					
Wheat	-0.700	0.083	0.226	0.121	-0.314
Rice	0.075	-0.650	-0.059	0.265	0.423
Sugar	0.164	-0.132	-0.462	-0.083	-0.294
Soyoil	0.034	0.073	0.006	-0.736	-0.202
Cornmeal	-0.126	0.133	-0.082	-0.386	-0.613
Quartile 2					
Wheat	-0.668	0.113	0.266	0.146	-0.013
Rice	0.096	-0.630	-0.033	0.283	0.607
Sugar	0.207	-0.092	-0.416	-0.040	0.005
Soyoil	0.038	0.077	0.014	-0.736	-0.134
Cornmeal	-0.087	0.169	-0.044	-0.342	-0.391
Quartile 3					
Wheat	-0.615	0.128	0.295	0.177	-0.008
Rice	0.122	-0.629	-0.019	0.286	0.613
Sugar	0.253	-0.081	-0.391	-0.018	0.011
Soyoil	0.048	0.075	0.019	-0.737	-0.131
Cornmeal	-0.055	0.176	-0.027	-0.327	-0.387
Quartile 4					
Wheat	-0.574	0.150	0.305	0.230	-0.016
Rice	0.146	-0.622	-0.015	0.310	0.609
Sugar	0.288	-0.068	-0.384	0.019	0.005
Soyoil	0.059	0.080	0.021	-0.725	-0.133
Cornmeal	-0.031	0.185	-0.022	-0.302	-0.391

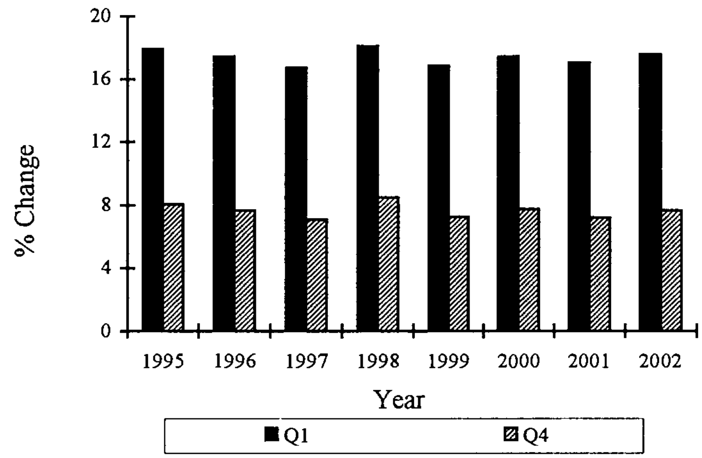

Figure 1. Percentage Change in Per Capita Chicken Consumption for Different Income Groups

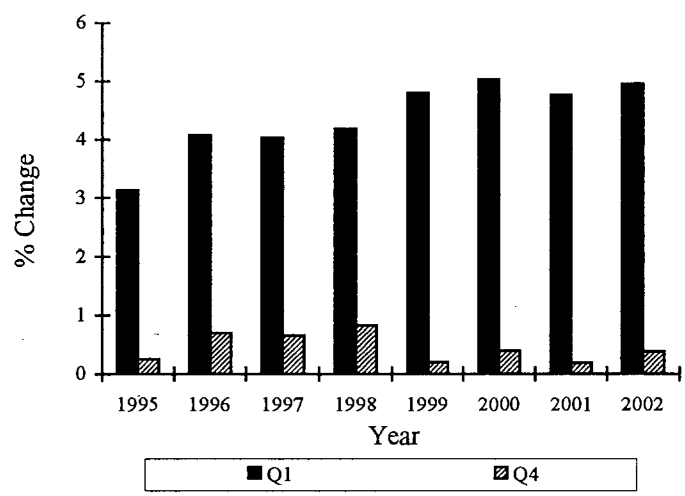

Figure 2. Percentage Change in Per Capita Wheat Consumption for Different Income Groups

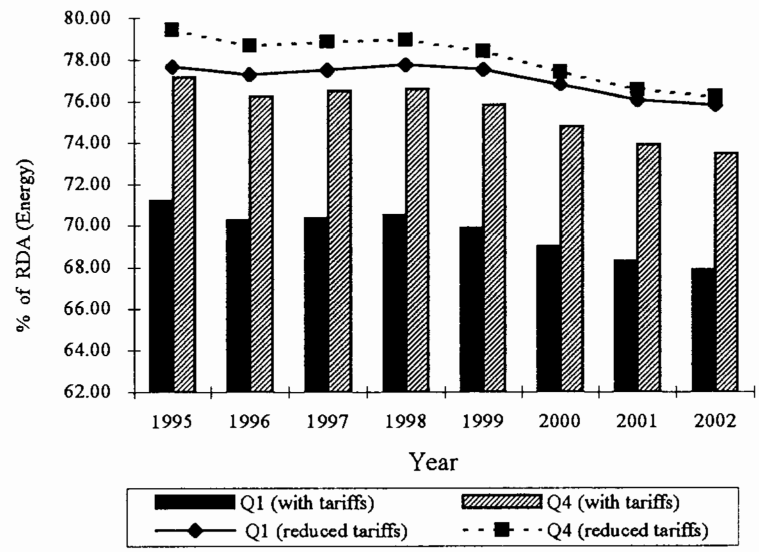

Figure 3. Percentage of Recommended Level of Energy Intakes for Different Income Groups

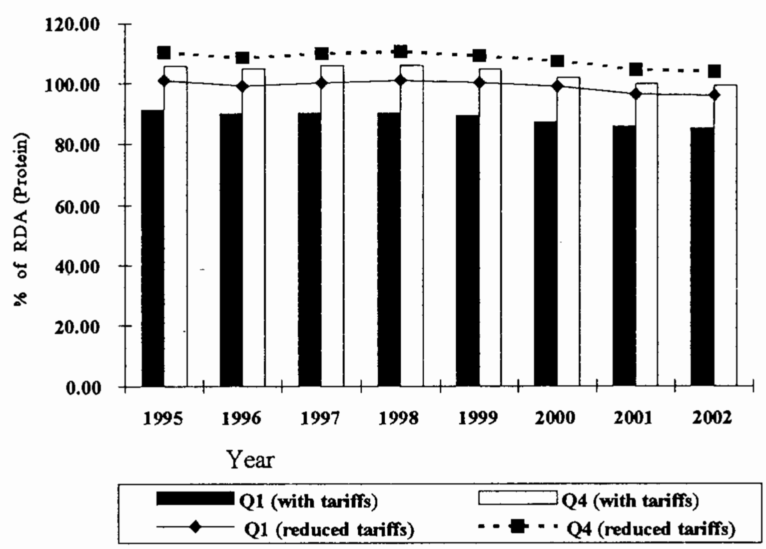

Figure 4. Percentage of Recommended Level of Protein Intakes for Different Income Groups

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