The Structure of Food Demand in Mexico: An Application of the Amemiya-Tobin Approach to the Estimation of a Censored System

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In contrast to demand analyses utilizing aggregated time-series data, the modeling of micro-level demand patterns necessitates the analyst explicitly incorporate household heterogeneity. In addition, the researcher needs to address the econometric problem of commodity purchase censoring especially when detailed commodity definitions are used. Although this issue has been well addressed in the analysis of expenditures on a single commodity, as evidenced by the large number of studies incorporating infrequency of purchase, double-hurdle and tobit specifications, the extension of these single equation models to demand systems estimation is more complicated. These complications arise out of the need for an empirical model that insures non-negativity of predicted quantities, incorporates constraints implied by economic theory, and the numerical problem of having to evaluate relatively high dimension cumulative density functions during model estimation.

Two approaches used to estimate micro-level demand systems include (i) the Kuhn-Tucker approach proposed by Wales and Woodland (1983) and associated dual suggested by Lee and Pitt (1986) and (ii) the Amemiya-Tobin approach proposed by Amemiya (1974) and operationalzed by Wales and Woodland (1983). The Kuhn-Tucker approach derives demand (share) equations from maximizing an explicitly specified random utility function after incorporating non-negativity and budget constraints. Lee and Pitt (1986)'s dual approach derives the demand (share) equations using Roy's Identity from a random indirect utility function and assumes that consumers compare virtual (reservation) prices to actual market prices in making their purchase decisions.

The main issue that must be addressed when using the Kuhn-Tucker approach and its dual, is the derivation of an estimable demand system. For some system specifications, it is not an easy task to specify a direct or indirect utility function that allows for system estimation. Furthermore, the coherency problem must be addressed where incoherency is characterized by the sum of purchase regime probabilities not equaling one. As noted by van Soest and Kapteyn (1993), van Soest and Kooreman(1990) and by Ransom(1987), an incoherent system will lead to inconsistent parameter estimates.

In contrast to the above, under the Amemiya-Tobin approach to the estimation of a censored demand system, demand (share) equations are derived from a non-stochastic utility function and latent expenditures (shares) are hypothesized to differ from observed expenditures due to (i) errors of maximization by the consumer, (ii) errors of measurement of the observed shares or (iii) other random disturbances which influence the consumer's decisions (Wales and Woodland, 1983). To account for these differences, error terms are added to the deterministic shares. Given the assumed normality of equation error terms, observed expenditures (shares) are thus normally distributed about the deterministic expenditures (shares). Non-negativity constraints are incorporated via the truncation of the above equation error terms similar to the censored multivariate Tobit model proposed by Amemiya(1974). Unlike the Kuhn-Tucker based approaches, incoherency is not a problem under this approach.

In this paper we extend the Amemiya-Tobin approach to demand systems estimation to an analysis of Mexican household food demand via the use of a linearly approximated Almost-Ideal Demand System (LA/AIDS) specification and a 1998 household expenditure survey. The LA/AIDS specification used here incorporates both nonnegativity and budget constraints. Given the disaggregated definition of commodities used in this analysis which requires we evaluate a high dimension truncated distribution of random variables, we use a simulated maximum likelihood (SML) technique to obtain parameter estimates. From this analysis the impact of changes in price, income and household characteristics on food demand can be quantified.

An estimable censored LA/AIDS model imposing adding-up and other theoretical constraints is developed in the following section where we develop the associated likelihood function which is transformed so that in can be used within our SML algorithm. We then develop, from a theoretical perspective, predictions of conditional and unconditional commodity shares. From these, associated elasticity measures are derived. We then present our empirical application of the analysis of Mexican household food demand. The final section provides some insights into future research efforts.

Censored LA/AIDS Demand Systems

Following Deaton and Muellbauer (1980), and Heien and Wessells (1990), we assume the consumer's utility function can be represented by a PIGLOG class from which the AIDS demand system is derived (Pollack and Wales, 1992). The following system of M+1 latent share (W^*) equations can be expressed as:

(1) $$W^* = U + \varepsilon,$$

where $U = A + \gamma \ln P + \eta \ln Y$, $A = \alpha + \beta X$, $Y = \frac{y^*}{P^*}$, P is a [M + 1] column vector of commodity prices, X is a [L x 1] vector of demographic characteristics, P^* is Stone's price index, y^* is a [(M + 1) x 1] vector of total expenditures, ε is a [(M + 1) x 1] vector of equation error terms.[1] Equation parameters are: α [(M+1) x 1], β [(M + 1) x L], γ [(M+1) x (M+1)], and η [(M+1) x 1].

Theoretical constraints such as homogeneity and symmetry can be imposed on (1). Notice however there are no non-negativity constraints imposed on these latent shares. There is nothing in the formulation to ensure that the elements of W^* lie between 0 and 1.

Given the budget constraint we know the latent shares must sum to one, the joint density function of ε is singular; one of the [M + 1] latent share equations must be dropped during estimation. Dropping the last equation from the estimation, we assume the first M share equations' error terms, ε in (1), are distributed multivariate normal with a joint probability density function (PDF). That is, $\varepsilon \sim N(0, \Sigma)$, where Σ is the [M x M] error variance-covariance matrix.

The mapping of the vector of latent, W^*, to observed shares, W, must take into account that elements of W: (i) lie between 0 and 1, and (ii) sum to unity for a particular observation. Following Wales and Woodland (1983), the following mapping rule imposes these characteristics (and omitting household subscripts):

$$(2) \qquad W_i = \begin{cases} \dfrac{W_i^*}{\sum_{j \in S} W_j^*}, & \text{if } W_i^* > 0, \\ 0, & \text{if } W_i^* \le 0, \end{cases}$$

where S is a set of all positive shares' subscripts. As pointed out by Wales and Woodland (1983), though there may be ways other than (2) in mapping W^* to W, the one we have chosen is both simple and has the property that the resulting density function is independent of whatever set of the W^*'s is used in its derivation.

Assuming that at least one commodity is purchased, we can partition observed purchase patterns into three general purchase regimes: (i) at least one commodity is purchased but the total number of purchased commodities is less then M, (ii) M commodities are purchased, and (iii) all M +1 commodities are purchased. For each of these regimes we can develop regime-specific likelihood functions that can be used to obtain system parameter estimates. Since a particular household is associated with only one purchase regime, the likelihood function appropriate for its purchase pattern determines the contribution this household makes to the overall sample likelihood function value.

Derivation of the Likelihood Function for Regime I: Some Commodities Not Purchased

For households where K commodities are purchased and $M > K \geq 1$, we can rearrange the ordering of the M commodities so that the first K are purchased. The error term covariance matrix, Σ, can be partitioned as:

(3) $$\Sigma = \begin{bmatrix} \Sigma_{11} & \Sigma_{10} \\ \Sigma_{10}' & \Sigma_{00} \end{bmatrix},$$

where Σ_{11} is a K x K submatrix associated with the purchased commodities, Σ_{00} is a (M-K) x (M-K) submatrix associated with the non-purchased commodities, and Σ_{10} is a (M-K) x K submatrix of covariance across purchased and non-purchased commodities. With this rearrangement, the likelihood of a household being in a purchase regime where the first K commodities are positive and zero for the remaining can be represented via the following (Wales and Woodland, 1983):

$$(4)\quad L\left(W_1, W_2, \cdots, W_k > 0;\; W_{k+1} = W_{k+2} \cdots = W_M = 0\right)$$

$$= \int_{W_1}^{+\infty} \int_{1-\frac{W_1^*}{W_1}}^{0} \int_{1-\frac{W_1^*}{W_1}-W_{k+1}^*}^{0} \cdots \int_{1-\frac{W_1^*}{W_1}-W_{k+1}^*-\cdots-W_M^*}^{0} \phi\left(W_1^*, W_2^*, \cdots, W_M^*; U, \Sigma\right) dW_M^* \cdots dW_{k+1}^* dW_1^* \cdot$$

The integral in (4) is [M–K+1] fold, which is the number of non-purchased commodities plus one. In order to evaluate the multivariate integrals, as we will discuss below, we transform equation (4) as follows by reducing the dimension of $f(\cdot)$ from M to [M–K+1]:

$$(5)\quad L\left(W_1, W_2, \cdots, W_k > 0;\; W_{k+1} = W_{k+2} \cdots = W_M = 0\right)$$

$$= B \cdot \int_{W_1}^{+\infty} \int_{1-\frac{W_1^*}{W_1}}^{0} \int_{1-\frac{W_1^*}{W_1}-W_{k+1}^*}^{0} \cdots \int_{1-\frac{W_1^*}{W_1}-W_{k+1}^*-\cdots-W_M^*}^{0} \phi\left(W_1^*, W_{k+1}^*, \cdots, W_M^*; U^*, \Omega_{11}\right) dW_M^* \cdots dW_{k+1}^* dW_1^*,$$

where $U^* = \begin{pmatrix} U_1^* \\ U_{k+1}^* \\ \vdots \\ U_M^* \end{pmatrix} = \Omega_{11}\Omega_{10}^{-1} \begin{pmatrix} U_1 \\ U_{k+1} \\ \vdots \\ U_M \end{pmatrix}$, an [(M-K+1) x 1] vector, and

$$B = (2\pi)^{\frac{1-k}{2}} \cdot |\Sigma|^{-\frac{1}{2}} \cdot |\Omega_{11}|^{\frac{1}{2}} \cdot e^{-\frac{1}{2}\left\{ \begin{pmatrix} U_1 \\ U_{k+1} \\ \vdots \\ U_M \end{pmatrix}' \Omega_{00}^{-1} \begin{pmatrix} U_1 \\ U_{k+1} \\ \vdots \\ U_M \end{pmatrix} - \begin{pmatrix} U_1^* \\ U_{k+1}^* \\ \vdots \\ U_M^* \end{pmatrix}' \Omega_{11}^{-1} \begin{pmatrix} U_1^* \\ U_{k+1}^* \\ \vdots \\ U_M^* \end{pmatrix} \right\}}.$$

The above Ω_{ij}'s are [(M-K+1) x (M-K+1)] matrices defined as:

$$\Omega_{11} = \begin{bmatrix} I_K'\sigma_{11}I_K & I_K'\sigma_{10} \\ \sigma_{10}'I_K & \sigma_{00} \end{bmatrix},\ \Omega_{00} = \begin{bmatrix} J'\sigma_{11}J & J'\sigma_{10} \\ \sigma_{10}'J & \sigma_{00} \end{bmatrix}, \text{ and } \Omega_{10} = \begin{bmatrix} I_K'\sigma_{11}J & I_K'\sigma_{10} \\ \sigma_{10}'J & \sigma_{00} \end{bmatrix}.$$

where I_K is a [K x 1] vector of ones, and J is a [K x 1] vector with elements:

$\left(1, \frac{U_2}{\left(\frac{W_2}{W_1}\right)U_1}, \frac{U_3}{\left(\frac{W_3}{W_1}\right)U_1}, \frac{U_4}{\left(\frac{W_4}{W_1}\right)U_1}, \cdots, \frac{U_k}{\left(\frac{W_k}{W_1}\right)U_1}\right)'$. The σ_{ij}'s are defined via the following [M x

M] matrix: $\begin{bmatrix} \sigma_{11} & \sigma_{10} \\ \sigma_{10}' & \sigma_{00} \end{bmatrix} = \begin{bmatrix} A\,\Sigma_{11}^{-1}\,A' & A\,\Sigma_{10}^{-1} \\ \Sigma_{10}^{-1'}\,A' & \Sigma_{00}^{-1} \end{bmatrix}^{-1}$. The [K x K] matrix *A* is a diagonal matrix with elements: $\left(1, \frac{W_2}{W_1}, \frac{W_3}{W_1}, \cdots, \frac{W_k}{W_1}\right)$. Finally, the Σ_{ij}^{-1} matrices are obtained from the full error variance matrix, Σ, in (3).

From the results shown in Tallis (1965), the likelihood function represented by (5) can be further transformed to:

(6) $L\left(W_1, W_2, \cdots, W_k > 0;\; W_{k+1} = W_{k+2} \cdots = W_{M+1} = 0\right) = B \cdot \Phi_{M-k+1}\left(b; R_C\right)$,

where $\Phi_{M-k+1}\left(b; R_C\right)$ is a [M-K+1] dimensional multivariate standard normal cdf with correlation coefficient matrix as R_C and evaluated at vector b. Vector b is [(M-K+1) x 1] and can be shown to be equal $E \cdot P$, where E is a [M-K+1] diagonal matrix with nonzero elements: $\left(\left(C_1 R C_1'\right)^{-1/2}, \left(C_{k+1} R C_{k+1}'\right)^{-1/2}, \cdots, \left(C_M R C_M'\right)^{-1/2}\right)$. Where

$$C = \begin{pmatrix} C_1 \\ C_{k+1} \\ \vdots \\ C_M \end{pmatrix} = H \cdot D^{\frac{1}{2}},\; H = \begin{bmatrix} \frac{1}{W_1} & 1 & 1 & \cdots & 1 \\ 0 & -1 & 0 & \cdots & 0 \\ 0 & 0 & -1 & \cdots & 0 \\ \vdots & \vdots & \vdots & \cdots & \vdots \\ 0 & 0 & 0 & \cdots & -1 \end{bmatrix}$$

, a [M-K+1] square matrix, R is the correlation coefficient matrix derived from Ω_{11}, and D the diagonal elements of Ω_{11}.

Term $P = \begin{pmatrix} 1 - H_1 U_1^* \\ -U_{k+1}^* \\ \vdots \\ -U_M^* \end{pmatrix}$, where H_1 is the first row of matrix H. The new correlation coefficient matrix (R_C) is given as $R_C = ECRC'E'$ (Tallis, 1965).

Equation (6) represents a rectangular standard multivariate normal probability, which can be conveniently evaluated using standard simulation procedures. The smooth recursive conditioning simulator (GHK) suggested by Geweke(1991), Hajivasiliou and McFadden(1990), and Keane(1993,1994) is adopted for this analysis to simulate this multivariate normal probability.

Derivation of the Likelihood Function for Regime II: Only One Commodity Not Purchased

Regime II is characterized by the number of commodities actually purchased, K, equaling M. This implies that (5) can be restated as:

$$(7)\quad L\left(W_1, W_2, \cdots, W_M > 0; W_{M+1} = 0\right) = B \cdot \int_{W_1}^{+\infty} \phi\left(W_1^*; U^*, \Omega_{11}\right) dW_1^*$$

where $U^* = U_1^* = \Omega_{11}\Omega_{10}^{-1}U_1$, and $\Omega_{11} = I'_M \sigma_{11} I_M$, $\Omega_{00} = J'\sigma_{11}J$, $\Omega_{10} = I'_M \sigma_{11} J$, are all scalars now with $\sigma_{11} = \left(A\,\Sigma^{-1}\,A'\right)^{-1}$, where A is an (M x M) diagonal matrix with

elements: $\left(1, \frac{W_2}{W_1}, \frac{W_3}{W_1}, \cdots, \frac{W_M}{W_1}\right)$, $J = \left(1, \frac{U_2}{\left(\frac{W_2}{W_1}\right)U_1}, \frac{U_3}{\left(\frac{W_3}{W_1}\right)U_1}, \frac{U_4}{\left(\frac{W_4}{W_1}\right)U_1}, \cdots, \frac{U_M}{\left(\frac{W_M}{W_1}\right)U_1}\right)'$, and

$B = (2\pi)^{\frac{1-k}{2}} \cdot |\Sigma|^{-\frac{1}{2}} \cdot |\Omega_{11}|^{\frac{1}{2}} \cdot e^{-\frac{1}{2}\left\{U_1'\Omega_{00}^{-1}U_1 - U_1^{*'}\Omega_{11}^{-1}U_1^*\right\}}$. Thus, under Regime II, the likelihood function requires the integration of a univariate PDF.

Derivation of the Likelihood Function for Regime III: All Commodities Purchased

For households where all commodities are purchased (K = M+1), the likelihood function of this regime is just the [M x 1] multivariate PDF of error term, ε, which is defined in (1) and distributed as $MN(0,\Sigma)$, where Σ is given by (3). That is:

(8) $L\left(W_1, W_2, \cdots, W_{M+1} > 0\right) = \phi(\varepsilon)$

Consistent and efficient estimates of parameters can be obtained by maximizing the sum of log likelihood function over all households, which fall into one of the three demand regimes, i.e., equations (6) - (8).

Evaluation of Predicted Shares and Demand Elasticities

Expected values of observed expenditure shares can be obtained from our censored demand system by summing the products of each regimes probability and expected conditional share values over all possible regimes. Let R_k represent the k^{th} demand regime that is characterized as:

$$R_k = \left(W_1 = W_2 = \cdots = W_k = 0; W_{k+1} > 0, \cdots, W_{M+1} > 0\right).$$

The expected value of the j^{th} observed share is:

(9) $$E\left(W_j\right) = \sum_{k=1}^{M+1} \alpha_{R_k} E\left(W_j \mid R_k\right),$$

where α_{R_k} is the probability of regime R_k occurring, and

(10) $$\alpha_{R_k} = \text{prob}\left(R_k\right) = \text{prob}\left(W_1 = W_2 = \cdots = W_k = 0; W_{k+1} > 0, \cdots, W_{M+1} > 0\right)$$

$$= \int_{-\infty}^{-U_1} d\varepsilon_1 \int_{-\infty}^{-U_2} d\varepsilon_2 \cdots \int_{-\infty}^{-U_k} d\varepsilon_k \int_{-U_{k+1}}^{\sum_{i=k+2}^{M+1} U_i - \sum_{i=2}^{k} \varepsilon_i} d\varepsilon_{k+1} \cdots \int_{-U_{M-1}}^{\sum_{i=M}^{M+1} U_i - \sum_{i=2}^{M-2} \varepsilon_i} d\varepsilon_{M-1} \int_{-U_M}^{U_{M+1} - \sum_{i=2}^{M-1} \varepsilon_i} \phi\left(\varepsilon_1, \varepsilon_2, \cdots, \varepsilon_M\right) d\varepsilon_M,$$

where $\phi\left(\varepsilon_1, \varepsilon_2, \cdots, \varepsilon_M\right)$ is the multivariate normal pdf with mean vector and variance-covariance as given in equation (2). The expected share value conditional on purchase regime R_k can be represented as:

(11) $$E\left(W_j \mid R_k\right)=\begin{cases}\dfrac{E(W_j^* \mid R_k)}{\sum\limits_{i=k+1}^{M+1} E(W_i^* \mid R_k)}, & \text{if} \quad j>k,\\ 0, & \text{otherwise}\end{cases}$$

with $E\left(W_j^* \mid R_k\right)=U_j+E\left(\varepsilon_j \mid R_k\right)=U_j+\dfrac{\alpha_{R_k}^{\varepsilon_j}}{\alpha_{R_k}}$, where,

(12) $$\alpha_{R_k}^{\varepsilon_j}=\int_{-\infty}^{-U_1} d\varepsilon_1 \int_{-\infty}^{-U_2} d\varepsilon_2 \cdots \int_{-\infty}^{-U_k} d\varepsilon_k \int_{-U_{k+1}}^{\sum\limits_{i=k+2}^{M+1} U_i-\sum\limits_{i=2}^{k}\varepsilon_i} d\varepsilon_{k+1} \cdots \int_{-U_{M-1}}^{\sum\limits_{i=M}^{M+1} U_i-\sum\limits_{i=2}^{M-2}\varepsilon_i} d\varepsilon_{M-1} \int_{-U_M}^{U_{M+1}-\sum\limits_{i=2}^{M-1}\varepsilon_i} \varepsilon_j \phi\left(\varepsilon_1, \varepsilon_2, \cdots, \varepsilon_M\right) d\varepsilon_M$$

From (9) the impact of changes in prices, demographics or expenditures on food demand can be obtained but one needs to evaluate M-dimension integrals shown in (10) and (12). Given that there are 2^{M+1}-1 purchase regimes, one may need to evaluate (10) and (12) a large number of times for a reasonably sized demand system. Phaneuf, Kling, and Herriges (2000) in an analysis of a censored demand system applied to recreation choices, develop a simulation procedure to evaluate expressions similar to (9)-(12). We modify their procedure to our application.

Assume we have R replicates of the [M+1] error term vector, ε in (1). The r^{th} simulated latent share, $E\left(W^*\right)_r$, evaluated at sample means of our exogenous variables (indicated by a bar over a variable) is

(13) $$E\left(W^*\right)_r=\alpha+\gamma \ln \overline{P}+\beta \ln \frac{\overline{Y}}{\overline{\overline{P^*}}}+\varepsilon_r$$

where ε_r is the r^{th} replicate of ε. The r^{th} replicate of the i^{th} observed share then is

$$(14) \qquad E(W_i)_r = \begin{cases} \dfrac{E(W_i^*)_r}{\sum_{j \in S} E(W_j^*)_r}, & \text{if } E(W_i^*)_r > 0, \\ 0, & \text{if } E(W_i^*)_r \le 0, \end{cases}$$

where subscript i represents the i^{th} element of W. The expected observed share vector for R replicates is then calculated as simple average of these simulated values:

$$(15) \qquad E(W) = \frac{1}{R} \sum_{r=1}^{R} E(W)_r \,.$$

Suppose we have a small change in price j, ΔP_j, the elasticity vector with respect to this price change is:

$$(16) \qquad \eta_j = -\delta_j + \frac{\Delta E(W)}{\Delta P_j} \cdot \frac{P + \frac{\Delta P_j}{2}}{E(W) + \frac{\Delta E(W)}{2}},$$

where δ_j is a vector of 0's with the j^{th} element 1, and $\Delta E(W)$ is the change of the simulated E(W) given the change of price, ΔP_j.

The Analysis of Food Purchases in Mexico

In this empirical application, we estimate a censored food demand system for Mexican households using the above regime specific likelihood functions. Mexico represents a significant export market for raw and processed U.S. food products, and recently become the U.S.'s third largest trading partner after the European Union and Canada. The study of the effects on Mexican household food purchase patterns will provide valuable information for food marketing managers crafting export policies.

Description of the Mexican Household Survey Data

The data used in this analysis is the 1998 *Encueta Nacional de Ingresos y Gastos de Los Hogares (ENIGH)*, a nationwide survey of Mexican household food and non-food purchases, household cash and non-cash income and other household socioeconomic characteristics. This survey was undertaken by the Instituto Nacional de Estadistica, Geografia e Informatica over the August-November 1998 period. Surveyed households maintained weekly diaries of their daily expenses. The survey data contained not only purchase information, but also a detailed set of household and member-specific information. To avoid problems with respect to the valuation of home produced goods, we limited our current analysis to urban residents. That is we only included households that resided in towns with greater than 15,000 persons. From this subset we then randomly selected 50% of the households due to the complexity of the econometric model described above. We also excluded households that did not record any expenditures on food for at-home consumption during the survey week. Our final sample size was 2,972 households.

Table 1 presents purchase frequencies, means, and standard deviations of conditional expenditures for food categories used in our system estimation. Table 2 provides the definition of exogenous variables (excluding unit values) used in the LA-AIDS model (equation 1) along with sample means and standard deviations. Included in the analysis are measures of household size and age composition, a variable indicating ownership of a refrigerator/freezer and a set of regional dummy variables.

Commodity prices are not explicitly contained in the ENIGH survey. Instead we used calculated unit-values as measures of price given that we have commodity expenditures as well as purchased amounts in the data set. For households not purchasing a particular commodity, we adopt a zero-order correction procedure where unobserved unit-values are replaced with the average unit-value obtained by purchasing households in the same area, represented by state of residence and degree of urbanization. Given the complexity of the model the endogenization of product quality and unit-values that was addressed in the single-equation meat demand analysis by Dong and Gould (2000) was not undertaken in this version of our model. An extension of our demand system would allow one to account for this endogenization.

Summary of Estimated Demand System Coefficients

Given the use of 12 commodity categories, and 13 demographic variables, a total of 297 parameters were estimated using the GAUSS software system and BHHH maximum likelihood procedure. Table 3 shows the maximum likelihood parameter estimates for the demographic, expenditure and unit-value related coefficients. The equation omitted during estimation was the one corresponding to fluid milk. The associated parameters for this omitted equation are retrieved from the LA-AIDS adding-up, symmetry, and homogeneity constraints.

Of the 156 demographic related parameters estimated, 106 (68%) were statistically significant at the 0.10 level of significance and 93 (60%) were significant at the 0.05 level. There is evidence of significant differences in food purchase patterns across regions as well as significant impacts of household composition and refrigerator ownership.

In addition to the demographic related parameters, Table 3 also shows the estimated own and cross-price coefficients. All of the own-price coefficients were found to be statistically different from zero at the 0.05 level of significance. Of the 66 cross-price coefficients estimated, 23 were statistically significant at the 0.10 level (35%) and of these, all but 2 were significant at the 0.05 level (32%). Of the 12 estimated expenditure coefficients, 8 were statistically significant at the 0.10 level and all but one of these at the 0.05 level.

Estimated Price and Demographic Elasticities

The estimated parameters themselves are of little interest. From these parameters however we estimate uncompensated, unconditional own and cross-price elasticities by evaluating equations (9)-(12) and using the simulation procedures outlined by Phaneuf, Kling, and Herriges (2000).[2] The resulting elasticity estimates are shown in Table 4. As expected, all own-price elasticities were found to be negative with a range of -0.133 for raw (unprocessed) pork products to –1.1425 for fluid milk.

Golan, Perloff and Shen (2001) use a maximum entropy approach to examine the structure of Mexican meat purchases using a 1992 version of the ENIGH data set. Although there were significant changes in the Mexican economy between 1992 and 1998 (e.g., peso devaluation in 1994), comparing our price elasticities provides an indication of their reasonableness. Table 5 provides a comparison of the price elasticities of the above analysis with our current study. Several major differences are evident. First, the own price elasticity of beef are substantially larger in the previous analysis, Second, the own and cross-price effects of seafood price changes under the previous analysis are much larger than the present analysis. The degree to which this is due to the differing demand system estimation is unclear. Third, under the current analysis, pork and beef were estimated to be gross complements compared to substitutes under the earlier study.[3] In contrast, we estimate that poultry and beef to be substitute products which is in contrast to Golan, Perloff and Shen (2001).

Table 6 shows estimated demographic elasticities for the continuous demographic variables used in our analysis along with the percentage point change in shares due to a discrete change in the set of dichotomous demographic characteristics.[4] Surprisingly we find that the elasticity impact of a change in household size negatively impacts milk demand. In contrast, there is a positive impact on milk demand of having small children in the household. The impact of having refrigerated storage varies across commodity. Having refrigerated storage increases meant and fish demand. There is a decrease in reliance on grain-based products and dried beans. Demand for dairy products is also positively impacted by the presence of refrigerated storage in the home.

Future Research Directions

In this paper, we developed an estimable household demand system using an adapted Amemiya-Tobin approach to account for the censoring of commodity purchases. The model is estimated using simulated maximum likelihood techniques. The use of this technique has enabled use to evaluate a large (12 commodity) demand system which would have been impractical under traditional maximum likelihood techniques.

This research represents a first attempt at estimating a disaggregated food demand system. There are obvious changes that can be undertaken that could improve the quality of this research. First, given the debate concerning the use of Stone's Index in the LA-AIDS model a revised version of this model will incorporate an appropriate modification of the traditional index (Moschini. 1995, Buse and Chan, 2000). Second, given the inherent non-linearity of our estimation process, the use of the non-linear AIDS specification may want to be attempted. A third methodological improvement to the current specification would be the endogenization of product quality similar to the procedures outlined in the single equation approaches of Dong, Shonkwiler and Capps(1999) and Dong and Gould (2000). That is, in spite of the estimation of a disaggregated demand system, there continues to be a range of product quality within each commodity group where this product quality is an endogenous variable that is part of the household's purchase experience.

Finally, the above analysis has quantified the unconditional impacts of the change in unit-values and household demographic characteristics on commodity demand. We need to analytically derive how the conditional demand levels and the probability of purchase are impacted by these changes. This will enable us to quantify both the intensive and extensive consumer response to changes in these variables.

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Table 1. Overview of Mexican Household Food Purchases

Commodity	% Households Purchasing	Mean Expenditure by Purchasing Households (Peso)	Std. Error of Conditional Expenditures (Peso)
Beans	51.3	15.6	21.4
Beef	69.2	46.6	51.9
Cheese	45.6	18.6	32.4
Fruits	62.4	21.9	33.9
Grains	97.9	47.0	33.9
Fluid Milk	80.5	35.5	38.7
Non-Alcoholic Bev.	75.8	30.5	34.3
Pork	25.7	30.9	69.3
Poultry	60.3	32.0	39.5
Processed Meat	55.1	23.8	36.9
Fish/Shellfish	22.1	26.3	71.3
Vegetables	86.8	29.8	27.2

Source: 2972 Randomly selected urban households from the 1998 ENIGH

Table 2. Demographic Characteristics Used in the Econometric Model

Variable	Description	Mean	Standard Deviation
REFRIG	Household owns a refrigerator/freezer (0/1)	81.9	-----
Household Size/Composition			
HHSIZE	Number of household members (#)	4.2	2.0
PERLT6	Percent of Household Members < 6 Years Old (%)	10.8	15.6
PER6_15	Percent of Household Members Between 6 and 15 Years Old (%)	18.5	20.3
PER16_24	Percent of Household Members Between 16 and 24 Years Old (%)	17.1	22.5
PERGT65	Percent of Household Members Aged 65 or More (%)	6.4	20.2
Region of Residence			
DF*	Distrito Federal, Estado de Mexico and Metropolitan Areas surrounding Mexico City	30.8	-----
NW	Baja California, Baja California Sur, Sonora and Sinaloa	8.9	-----
NE	Coahuila, Chihuahua, Nuevo Leon and Tamaulipas	12.0	-----
NC	Durango, San Luis Potosi, Queretaro and Zacatecas	6.5	-----
WEST	Nayarit, Jalisco, Colima, Guanajuato and Michocacan	17.6	-----
CENTRAL	Aguascalientes, Hidalgo, Morelos, Puebla and Tlaxcala	8.7	-----
SOUTH	Guerrero, Oaxaca and Veracruz	6.4	-----
SE	Yucatan, Tabasco, Quintana Roo, Chiapas and Campeche	9.1	

Note: *Indicates region used as the base region. In the econometric model we used the inverse of household size .

Table 3. Censored Demand System Parameter Estimates

	Beef	Pork	Poul	PrcMT	Seafood	Veg	Fruit	Grain	Beans	Cheese	NAB	Milk
Intercept	0.1117	-0.0080	0.1283	0.0549	-0.0004	0.1501	-0.0164	0.5852	0.0283	0.0239	-0.1416	0.084
Demographic Characteristics												
1/HHSIZE	-0.1010	-0.0422	-0.0882	0.0230	0.0006	-0.0521	0.0586	-0.1424	-0.0488	-0.0092	0.3023	0.0993
REFRIG	0.0560	-0.0008	0.0188	0.0264	0.0096	-0.0264	0.0364	-0.1305	-0.0567	0.0167	-0.0090	0.0594
PERLT6	-0.0516	-0.0404	-0.0495	0.0310	-0.0390	-0.0617	-0.0204	-0.0041	-0.0162	-0.0211	0.0633	0.2096
PER6_15	-0.0524	0.0013	-0.0448	0.0256	-0.0087	-0.0640	-0.0222	0.1036	0.0208	-0.0212	-0.0118	0.0738
PER16_24	-0.0059	0.0045	-0.0152	0.0114	-0.0047	-0.0317	-0.0259	0.0515	0.0125	-0.0051	-0.0037	0.0123
PERGT65	-0.0075	-0.0224	0.0284	-0.0505	-0.0200	0.0092	0.0041	0.0837	0.0440	0.0109	-0.1631	0.0832
NW	-0.0241	-0.0108	-0.0949	-0.0073	0.0305	-0.0639	-0.0570	0.0118	0.0145	0.0032	0.1299	0.0681
NE	0.0273	-0.0376	-0.0835	-0.0357	-0.0197	-0.0763	-0.0768	0.0581	0.0107	-0.0223	0.1730	0.0827
NC	-0.0302	-0.0388	-0.1169	-0.0411	-0.0196	-0.0310	-0.0341	0.0757	0.0313	0.0056	0.0986	0.1004
WEST	-0.0142	-0.0115	-0.0775	-0.0369	0.0006	-0.0341	-0.0340	0.0468	0.0324	-0.0108	0.0864	0.0528
CENTRAL	-0.0634	0.0157	-0.0465	-0.0121	-0.0183	0.0096	-0.0158	0.0736	0.0120	-0.0083	0.0125	0.0409
S	-0.0379	0.0175	-0.0289	-0.0216	0.0341	-0.0337	-0.0374	0.0568	0.0061	0.0127	0.0531	0.0208
SE	-0.0245	0.0765	0.0239	-0.0484	0.0154	-0.0576	-0.0514	0.0198	0.0223	-0.0482	0.1361	0.0639
Total Expenditures												
	-0.0126	-0.0114	-0.0078	-0.0022	-0.0063	0.0252	0.0323	-0.0424	0.0149	-0.0022	0.0468	-0.0342

(continued)

Table 3. Censored Demand System Parameter Estimates (Continued)

	Beef	Pork	Poul	PrcMt	Seafood	Veg	Fruit	Grain	Beans	Cheese	NAB	Milk
Beef	0.1025											
Pork	-0.0299	0.0796										
Poul	0.0257	-0.0069	0.0325									
PrcMt	-0.0155	-0.0153	-0.0039	0.0281								
Seafood	-0.0087	-0.0028	-0.0124	-0.0087	0.0284							
Veg	0.0036	-0.0013	0.0027	-0.0071	0.0073	0.0366						
Fruit	-0.0155	-0.0067	-0.0008	0.0089	-0.0058	-0.0169	0.0341					
Grain	-0.0375	-0.0165	-0.0258	0.0152	0.0111	-0.0288	0.0078	0.0782				
Beans	-0.0155	0.0162	-0.0015	-0.0131	-0.0041	-0.0048	-0.0105	-0.0284	0.0469			
Cheese	-0.0173	-0.0062	-0.0035	0.0087	-0.0075	-0.0085	0.0032	0.0063	-0.0017	0.0265		
NAB	0.0016	-0.0069	-0.0046	0.0016	-0.0023	0.0022	-0.0013	-0.0011	0.0032	0.0000	0.0100	
Milk	0.0063	-0.0033	-0.0015	0.0011	0.0055	0.0148	0.0034	0.0196	0.0132	0.0000	-0.0023	-0.0569

Note: The dark shaded cells indicate coefficients with ratio of the estimated coefficient and coefficient standard error exceed 2.0 . Light shaded cells indicate significance at between 1.64 and 2.00. Due to space constrains estimated coefficient values of the error term variance/covariance matrix can be obtained from the authors upon request.

Table 4. Simulated Expenditure and Uncompensated Price Elasticities

	Beef	Pork	Poul	PrcMt	Seafood	Veg	Fruit	Grain	Beans	Cheese	NAB	Milk
Beef	-0.5295	-0.1070	0.1244	-0.0712	-0.0344	0.0258	-0.0699	-0.1689	-0.0604	-0.0755	0.0099	0.0251
Pork	-0.2997	-0.1322	-0.0713	-0.1612	-0.0162	0.0007	-0.0723	-0.1581	0.2039	-0.0691	-0.0686	-0.0297
Poul	0.1696	-0.0251	-0.7795	-0.0289	-0.0758	0.0251	-0.0036	-0.1619	0.0026	-0.0186	-0.0286	-0.0138
PrcMt	-0.1369	-0.1084	-0.0352	-0.7621	-0.0695	-0.0590	0.0763	0.1298	-0.1020	0.0762	0.0131	0.0043
Seafood	-0.1019	-0.0013	-0.1494	-0.1061	-0.6935	0.1138	-0.0676	0.1461	-0.0236	-0.0935	-0.0193	0.0727
Veg	-0.0048	0.0077	0.0016	-0.0551	0.0464	-0.7941	-0.1089	-0.2099	-0.0242	-0.0539	-0.0051	0.0573
Fruit	-0.1800	-0.0499	-0.0319	0.0560	-0.0515	-0.1759	-0.7126	-0.0024	-0.0882	0.0230	-0.0441	-0.0188
Grain	-0.1147	-0.0281	-0.0728	0.0517	0.0383	-0.0791	0.0317	-0.7249	-0.0767	0.0243	0.0100	0.0796
Beans	-0.2215	0.2087	-0.0396	-0.1669	-0.0396	-0.0801	-0.1324	-0.3752	-0.4460	-0.0180	0.0154	0.1220
Cheese	-0.1986	-0.0564	-0.0363	0.0945	-0.0846	-0.0901	0.0427	0.0664	0.0037	-0.7037	0.0007	-0.0063
NAB	-0.0390	-0.0269	-0.0485	-0.0108	-0.0161	-0.0168	-0.0215	-0.0763	0.0206	-0.0071	-0.9756	-0.0421
Milk	0.0222	0.0144	0.0069	0.0011	0.0169	0.0503	0.0129	0.0530	0.0499	0.0038	0.0005	-1.1425
Tot.Exp.	0.9314	0.8738	0.9386	0.9735	0.9236	1.1431	1.2762	0.8606	1.1734	0.9681	1.2600	0.9108

Table 5. Comparison of Marshallian Price Elasticities,
Present Study and Estimates of Golan, Perloff and Shen (GPS,2000).

	Study	Beef	Pork	Poultry	Processed Meat	Fish/ Shellfish
Beef	Current	-0.5295	-0.107	0.1244	-0.0712	-0.0344
	GPS	-1.1023	0.0265	-0.1648	-0.1703	0.1060
Pork	Current	-0.2997	-0.1322	-0.0713	-0.1612	-0.0162
	GPS	0.1042	-0.5593	-0.2648	-0.2222	-0.2059
Poultry	Current	0.1696	-0.0251	-0.7795	-0.0289	-0.0758
	GPS	-0.0261	-0.0576	-0.6262	0.0052	-0.0403
Processed Meat	Current	-0.1369	-0.1084	-0.0352	-0.7621	-0.0695
	GPS	-0.1693	-0.1187	0.0919	-0.7830	0.4371
Fish/Shellfish	Current	-0.1019	-0.0013	-0.1494	-0.1061	-0.6935
	GPS	0.7522	-0.5534	-0.4093	1.1079	-2.1454

Table 6: Elasticity and Unconditional Predicted Share Impacts of Changes in Demographic Characteristics

Exogenous Variable	Beef	Pork	Poultry	Processed Meat	Fish/ Shellfish	Veg	Fruit	Grain	Beans	Cheese	NAB	Milk
	Elasticities											
HHSIZE	0.1445	0.1439	0.1782	-0.0670	-0.0103	0.0881	-0.1687	0.1434	0.1678	0.0221	-0.5141	-0.0642
PERLT6	-0.0247	-0.0476	-0.0336	0.0312	-0.052	-0.0374	-0.0167	-0.0022	-0.0186	-0.0242	0.0331	0.0543
PER6_15	-0.0454	0.0027	-0.0546	0.0425	-0.0196	-0.0703	-0.0349	0.0618	0.0451	-0.0457	-0.0159	0.0308
PER16_24	-0.0052	0.0081	-0.0175	0.0171	-0.0105	-0.0327	-0.0389	0.0282	0.0245	-0.0106	-0.0048	0.0035
PERGT65	-0.0020	-0.0157	0.0122	-0.0286	-0.0165	0.0032	0.0024	0.0169	0.0323	0.0077	-0.058	0.0145
	Change in Shares From Discrete Change in Dichotomous Exogenous Variable											
REFRIG	3.27	0.08	1.17	1.33	0.34	-1.19	1.81	-8.08	-2.43	0.74	-0.23	3.19
NW	-1.37	-0.33	-3.96	-0.38	0.96	-3.64	-2.4	0.4	0.61	0.12	7.84	2.16
NE	0.9	-1.05	-3.75	-1.7	-0.57	-4.52	-3.22	2.65	0.3	-0.89	9.88	1.97
NC	-1.93	-1.04	-4.71	-1.86	-0.55	-2.06	-1.6	3.99	1.25	0.12	5.31	3.08
WEST	-0.93	-0.36	-3.48	-1.66	0.01	-2.1	-1.56	2.68	1.38	-0.43	4.93	1.52
CENTRAL	-3.31	0.43	-2.17	-0.61	-0.5	0.48	-0.77	4.63	0.46	-0.35	0.47	1.24
S	-2.03	0.51	-1.38	-0.98	1.08	-2.03	-1.66	3.78	0.24	0.5	3.21	-1.24
SE	-1.56	2.62	1.02	-2.08	0.41	-3.46	-2.28	0.93	0.9	-1.64	8.41	-3.28

[1] As will be noted later, this preliminary version uses the commonly used Stone's Index in the LA-AIDS formulation. This can be easily modified to account for either alternative approximations such as the one suggested Moshchini(1994) or for estimating the nonlinear versions of the AIDS model.

[2] A future version of this analysis will allow us to calculate conditional elasticities.

[3] It should be noted that in the Golan, Perloff and Shen (2000) analysis, the Hicksian elasticity of a change on beef price on pork demand was not statistically significant.

[4] Except for the exogenous variable of concern, all exogenous variables are set at their mean values.