Household Size and Composition Impacts on Meat Demand in Mexico:

A Censored Demand System Approach

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Mexico represents a significant export market for raw and processed U.S. food products. With the adoption of the NAFTA, an expanding economy and growing population, Mexico has become the U.S.'s third largest trading partner after the European Union and Canada. Table 1 provides a summary of the role Mexico plays for U.S. agricultural exports. It's share of U.S. agricultural exports has steadily increased from less than 7% in 1990 to 12.7% in 2000. In 2001, U.S. food and agricultural exports to Mexico was $7.4 billion, an increase of nearly 58% since NAFTA was implemented in 1994. This trend is expected to continue given that by the end of 2003, nearly all Mexican import tariffs will be lifted. In addition, during April, 2002 the U.S. Secretary of Agriculture signed a joint agreement with her Mexican counterpart to create the Consultive Committee on Agriculture. This is a bi-lateral team with the mandate to strengthen the cooperation on agricultural trade issues between the two countries.

Given the anticipated increase in the importance of Mexico as an export market for U.S. agricultural products, it is important to understand the determinants of food purchase behaviour of Mexican households. To achieve this understanding we examine household food purchase patterns using a censored demand systems framework that allows for a disaggregated definition of foods. In contrast to single equation approaches such as that undertaken by Dong and Gould (2000), the demand system approach adopted here enables us to estimate both own and cross price elasticities, income elasticities and the effects of demographic characteristics or other variables that impact food demand.

Table 2 shows per capita food purchases for a 1998 urban sample of Mexican households.[1] Purchase amounts are obtained by dividing total household purchases by the number of household members. This count of household members is often used when defining per capita consumption or as a measure of household size. The implicit assumption associated with its use is that each household member has an equal impact on food purchases/expenditures. In reality, the impacts of household size will vary depending on the age/gender composition of household members (Deaton and Muellbauer, 1986).

One approach that can be used to avoid the assumption of equal expenditure impacts is to define household size via an endogenously determined *equivalence scale*. This scale can be used

to assign different weights to household members according to their age and gender (Deaton and Muellbauer, 1986).[2] Given the determination of an appropriate equivalence scale, a comparison of food expenditures for households of differing composition can be undertaken. As an example, suppose the weight given to a male adult between 25 and 45 years of age is 1.0, a female adult in the same age group a weight of .85 and a female child under 10 years of age a weight of .35, then a four-member household consisting of one male and two female adults and one female child in the above age groups would result in the household being composed of 3.05 *adult equivalents* (AE). A single parent household with one female adult would possess the corresponding adult equivalent of 1.20. The per capita expenditures patterns of these two households can then be compared where the number of AE are used as the expenditure deflator.

There are a number of approaches that have been suggested for the estimation of endogenously determined adult equivalent scales. These have ranged from the use of demographically translated utility consistent demand systems suggested by Barten (1964), Gorman (1976), Deaton and Muellbauer (1986) and implemented by Perali (1993) to single equation approaches used by Blokland(1976), Tedford, Capps and Havlicek (1986) and Muelbauer(1980).[3]

The present paper uses a demand system approach in the analysis of Mexican household meat purchases.[4] In this analysis and in contrast to Barten (1964) and Gorman (1976), we adopt a method where prices are scaled in such a manner that a single household food equivalent is estimated for each household instead of commodity specific scaling functions.[5] We limit ourselves to this one function given the number of additional parameters involved with commodity specific functions and the numerically intensive parameter estimation procedure described below.

As noted above, we limit our analysis to meat (including fish) purchases. From Table 2 we see that in 1998, 34% of total food-at-home expenditures are for meat or fish. Since large consumer surveys such as the one used here usually encompass short run purchases or consumption, zero values are commonplace. From Table 2 we see that the degree of censoring varies considerably across commodity. For example, less than 2% of household did not purchase grain-based products which include breads, tortillas, flours, etc. In contrast over 77% of surveyed households did not record a seafood purchase over the survey period. In the analysis to

follow we estimate a 5-equation meat demand system composed of beef, pork, poultry, processed meat, and fish/shellfish.[6]

Similar to single equation econometric models of food demand when there is significant censoring the use of standard estimation procedures that do not account for the non-negativity of such purchases are not appropriate. For this analysis we adopt the methodology originally proposed by Lee and Pitt (1986) which is based on a translog indirect utility function and the concept of virtual prices originally formulated by Neary and Roberts (1980) to account for the non-negativity of consumer demand.

Phipps(1998) estimated a non-censored translog demand system along with a theoretically consistent endogenous equivalence scale specification to examine the impact of children on food, clothing, transportation and housing expenditures. The analysis was limited to households where the only adults in the household are the male and female heads. Children where not differentiated by age or gender. Our analysis represents an extension of this original application not only via the utilization of a censored demand system but by differentiation of age composition of household members on meat expenditures. Given the number of commodities analyzed, our research also makes a significant contribution given use of simulated maximum likelihood techniques which do not require the imposition of restrictive distributional assumptions such as those used in Phaneuf et. al (2000).

Derivation of An Endogenously Determined Equivalence Scale

We assume that observed food purchase behavior can be obtained from a household's indirect utility function, V, which represents the maximum equally distributed equivalent indirect utility for each household member:

(1) $V = V(P, M \mid A) = Max \left[U(X \mid A, PX \leq M) \right]$

U represents a household's utility function, X a vector of consumed goods, A a vector of demographic characteristics, P a vector of market prices (unit values) faced by the household and M is total expenditure. That is, V represents the level of per capita utility which if it were shared by each household member would yield the same aggregate well-being as the actual distribution of utility within the household (Phipps, 1998). An equivalence scale, d(·), can then be defined using the above indirect utility function:

$$(2) \quad V = V(P, M \mid A) = V\left(P, \frac{M}{d} \mid A^R\right)$$

where A^R is the vector of characteristics of an arbitrarily defined reference household. Given (2), members of a household with characteristic vector A, facing prices P and household income M experience the same utility level as the reference household facing the same prices but with household income (M/d). As Blundell and Lewbel(1991) show, this equivalence scale can also be derived from the households' expenditure functions, $E(\cdot)$:

$$(3) \quad d = \frac{E(V, P \mid A)}{E(V, P \mid A^R)} = d(V, P \mid A)$$

Equivalence scales are of interest in that they enable the researcher to make inter-household comparisons of utilities and a determination of income levels at which members of households with different characteristics, such as member age or gender composition, are equally well off. If these equivalence scales are independent of utility level then preferences must satisfy *independence of base* (IB) and/or equivalence scale exactness (ESE).[7] Lewbel (1989) describes the general restrictions on cost and social welfare functions required for the estimation of IB equivalence scales. Blackorby and Donaldson (1993) show that to recover exact equivalence scales from demand behavior it is necessary that preferences not take a PIGLOG form.

As shown by (3), we need to specify a functional form for the equivalence scale measure. That is, we would like to define the equivalence of the reference household, V^R, such that

$$(4) \quad V(P, M \mid A) = V^R\left(P, \frac{M}{d(A, P)}\right)$$

We apply Roy's identity to the above indirect utility function to generate a system of demand equations. These demand equations will be functions of prices, income and demographic characteristics implying that the parameters of the equivalence scale can be obtained via the estimation of these demand equations (Blackorby and Donaldson, 1993).

In our analysis of meat expenditures we assume the household's indirect utility can be represented by the following nonhomothetic translog function:

$$(5) \quad \ln V(P,M) = \alpha_0 + \sum_{i=1}^{K} \alpha_i \ln(\tilde{p}_i) + \frac{1}{2}\sum_{i=1}^{K}\sum_{j=1}^{K} \beta_{ij} \ln(\tilde{p}_i)\ln(\tilde{p}_j) + \sum_{i=1}^{K} \varepsilon_i \ln(\tilde{p}_i)$$

where $\tilde{p}_i = p_i / M^*$, $M^* = \dfrac{M}{d(A,P)}$, $d(A,P) \equiv \prod_{s=1}^{S} N_s^{*\gamma_s} \exp\left(\sum_{l=1}^{L} A_l \Gamma_l\right) \prod_{i=1}^{K}\left(\dfrac{p_i}{M}\right)^{\left(\sum_{s=1}^{S} \delta_{is} N_s\right)}$,

K is the number of food commodities, p_i is the ith food's unit value (market price), S is the number of household member age classifications, L the number of demographic characteristics impacting household food expenditures (except for the number of household members), N_s the number of household members in the sth age classification other than that represented by the base household, $N_s^* = (N_s + 2)/2$, A_l is the lth demographic characteristic other than member category counts, and $\gamma_s's$, $\Gamma_l's, \alpha_i's, \beta_{ij}'s$, and $\delta_{is}'s$ are parameters to be estimated.[8] ε_i is an error term

where $\varepsilon \sim N(0,\Sigma)$ and Σ is the $(K \times K)$ error term covariance matrix. Suppose we define a reference household as a two-person household composed of a married couple between the age of 18 and 65. Except for the impact of demographic variables, A, the value of the scaling function is 1 for our reference household (e.g., $N_s = 0$ for each age group).

To insure symmetry, adding up, and homogeneity of degree zero in prices we impose the restrictions: $\beta_{ij} = \beta_{ji}$ and $\sum_{i=1}^{K} \alpha_i = -1$ and $\sum_{i=1}^{K} \delta_{is} = 0$ (s = 1,…,S) (Christenson, Jorgenson and Lau, 1975). It can be shown that this formulation satisfies general IB and ESE restrictions (Phipps, 1998).

From the above and via Roy's identity, we obtain the following share equations:

$$(6) \quad w_i \equiv \frac{p_i x_i}{M} = \frac{\alpha_i - \left(\sum_{s=1}^{S}\delta_{is}N_s\right) + \sum_{j=1}^{K}\beta_{ij}\ln(\tilde{p}_j) + \left(\sum_{s=1}^{S}\delta_{is}N_s\right)\sum_{j=1}^{K}\beta_j^*\ln(\tilde{p}_j) + \varepsilon_i}{-1 + \sum_{j=1}^{K}\beta_j^*\ln(\tilde{p}_j)} \quad (i=1,…,K)$$

$$= \frac{\left[\alpha_i - \left(\sum_{s=1}^{S} \delta_{is} N_s \right) + \sum_{j=1}^{K} \beta_{ij} \ln\left(\frac{p_j}{M} \right) + \beta_i^* \ln d(A,P) + \left(\sum_{s=1}^{S} \delta_{is} N_s \right) \sum_{j=1}^{K} \beta_j^* \ln\left(\frac{p_j}{M} \right) + \left(\sum_{s=1}^{S} \delta_{is} N_s \right) \sum_{j=1}^{K} \beta_j^* \ln d(A,P) + \varepsilon_i \right]}{-1 + \sum_{j=1}^{K} \beta_j^* \ln\left(\frac{p_j}{M} \right) + \sum_{j=1}^{K} \beta_j^* \ln d(A,P)}$$

where d(A,P) is defined via (5) and $\beta_i^* = \sum_{j=1}^{K} \beta_{ij}$, $\sum_{i=1}^{K} \alpha_i = -1$ and $\sum_{i=1}^{K} \delta_{is} = 0$.[9] With expenditure

shares summing to one, we can identify one of the error terms from the remaining and thus one share equation can be omitted from the estimation process.

In the analysis of non-censored commodity expenditures, Phipps(1998) uses the above to examine differences in household well-being when children are present in the household. Given the extended nature of Mexican households, we use this model as a base, but from (5) and (6) we formulate a more flexible model where we examine the impact on household food expenditures of the presence of household members in a series of age groups.

Estimation of A Censored Demand System

Consumption analyses based on time-series or aggregated-household data can reasonably incorporate the assumption that consumers respond to changes in prices, income, household composition, and other exogenous variables in a smooth continuous manner. In contrast, for disaggregated demand analyses such as being conducted here, the analyst needs to account for the distinct intensive and extensive consumption responses to changes in economic conditions. For example, with a drop in a commodity's price, current consumers of a normal good have an incentive to increase their consumption. This situation represents an intensive response, which has typically been analyzed with regression-based methodologies. For persons who are not current consumers of the commodity, a price reduction may induce them to enter the market and purchase the commodity, an extensive response. Given the discrete nature of the response to previous nonconsumers, and in contrast to the smooth adjustment process shown by current consumers, traditional regression methods may not be appropriate (Wales and Woodland , 1983, p. 263; Pudney, 1989, p. 138-39).

Within a single commodity framework, Heckman two-stage, Tobit, double-hurdle, and infrequency-of-purchase models are commonly used approaches to account for the above censoring of expenditures (Blundell and Meghir, 1987). In spite of accounting for purchase censoring within a systems framework being more numerically intensive from an econometrics perspective, increasing availability of simulated maximum likelihood methods has resulted in the application of these systems more common (Yen and Roe, 1989; Perali and Chavas, 2000; Phaneuf, 2000; Kao, Lee and Pitt, 2001;). These system approaches can be separated into two distinct types: those that do and do not explicitly incorporate a budget constraint.

Without a budget constraint, equations used to explain consumption of a separable commodity group can be treated as a group of correlated censored regressions (e.g. correlated Tobit equations). Pudney(1989) reviews the general framework of such models. Gould, Cornick and Cox(1994) apply such a system in an analysis of cheese purchases by U.S. households.

In contrast, Chiang and Lee(1992) develop a two-step procedure for estimating a random utility model that encompasses the discrete choice of whether or not to consume a particular commodity and the (nonnegatively) constrained quantity consumption decision. In this two-step procedure, a *multivariate* probability distribution incorporates the effect of censoring one commodity on other commodities in the system. Heien and Wessells (1990) in their household based analyses of food demand use *single-dimension* Heckman-type sample selection correction factors to control for the 0/1 purchase decision. Though attractive because of the ease with which their models can be estimated, correction factors obtained from *univariate* probit equations do not capture cross-commodity censoring impacts. As Shonkwiler and Yen (1999) and Vermuelen (2001) show, the above methodology is also inconsistent with the underlying theoretical model. They provide two-step alternatives that resulting in theoretically consistent results.

Wales and Woodland (1983) develop two approaches to modeling censored commodity demand based on both traditional Kuhn-Tucker conditions and those of Amemiya (1974). In their model, a direct utility function is maximized subject to budget and nonnegativity constraints. With the incorporation of these constraints, cross-equation restrictions must be placed on the demand (expenditure) functions and associated error covariance matrix.

Lee and Pitt (1986) formulate the dual to the Wales and Woodland (1983) approach where an indirect utility function is used to derive demand characteristics. Under their model,

consumers are assumed to compare virtual (reservation) prices (π) to actual market price (P) in making purchase decisions. Virtual prices represent the price level at which the consumer would be on the margin of consuming nonpurchased goods (Neary and Roberts, 1980; Pudney, 1989, p. 164-69). There are a number of analyses that have used either the primal or dual approaches for a variety of demand analyses (Gould, 1994; Phaneauf, 2000). For this analysis we adopt the Lee and Pitt (1986) framework for empirically implementing (6).

We assume an individual household maximizes utility, U(·), which is a continuously differentiable quasi-concave increasing function. Decision variables are consumption levels of N goods, x_i (i =1,...,N) chosen subject to a household's budget constraint. The consumer's problem can be represented by the maximization of the indirect utility function represented by (5). From this optimization process, x* is the optimal quantity vector, $x^* = \{0,\ldots,0, x^*_{m+1},\ldots, x^*_K\}$ where the first "m" commodities are not purchased. Virtual prices for these m commodities, π_i, and demand functions for the remaining (K-m) which can be shown to equal:

$$(7) \quad 0 = \frac{\partial V\left(\pi_1(P_O),\ldots,\pi_m(P_O)\mid P_O,\varepsilon\right)}{\partial \tilde{p}_i} \Rightarrow \pi_i(P_O) = v_i \frac{\partial U(x^*;\varepsilon)}{\partial x_i} \bigg/ \frac{\partial U(x^*;\varepsilon)}{\partial x_K} \quad (i=1,\ldots,m)$$

$$x_i = \frac{\dfrac{\partial V\left(\pi_1(P_O),\ldots,\pi_m(P_O)\mid P_O,\varepsilon\right)}{\partial \tilde{p}_i}}{\displaystyle\sum_{j=1}^{K}\tilde{p}\,\frac{\partial V\left(\pi_1(P_O),\ldots,\pi_m(P_O)\mid P_O,\varepsilon\right)}{\partial \tilde{p}_j}} \quad (i=m+1,\ldots,K)$$

where P_O is the set of market prices of the positively consumed goods.

The relative size of virtual and market prices determine whether a particular commodity is purchased. That is, the regime in which the first m of K commodities are not purchased is characterized by:

$$(8) \quad \pi_i(P_O) \le \tilde{p}_i \quad (i=1,\ldots,m)$$

With possible censoring of food purchases, (6) represents a set of latent share equations. Let C identify those commodities with zero expenditures, and O identify those that are purchased (e.g., W_C is the (m x 1) vector of zero-valued budget shares), P_O represent the ((K-m) x 1) vector of observed market unit values, π_C the (m x 1) vector of reservation prices for non-purchased goods and $\overline{P}_r = P_r/M$, (r=C,O). We can partition the symmetric price coefficient matrix according to the associated purchase regime:

$$\beta = \begin{bmatrix} \beta_{CC} & \beta_{CO} \\ \beta_{OC} & \beta_{OO} \end{bmatrix}; \ \beta_C \equiv \beta_{CC} 1'_m + \beta_{CO} 1'_{(K-m)}; \ \beta_O \equiv \beta_{OC} 1'_m + \beta_{OO} 1'_{(K-m)}.$$

Using the above share equations in (6) virtual prices for nonpurchased commodities are:[10]

$$(9) \quad w_C = 0 \rightarrow \alpha_C - \delta_C N + \beta_{CC} \ln(\pi_C) + \beta_{CO} \ln(\overline{P}_O) + \delta_C N \left[\beta'_C \ln(\pi_C) + \beta'_O \ln(\overline{P}_O) \right] +$$

$$\left[\gamma \ln(N^*) + \Gamma A + (\delta_C N)' \ln(\pi_C) + (\delta_O N)' \ln(\overline{P}_O) \right] \left[\beta_C + \delta_C N \left\{ 1'_C \beta_C 1_C \right\} \right] + \varepsilon_C = 0$$

$$\rightarrow \left[\beta_{CC} + \delta_C N \beta'_C + (\delta_C N)' \left[\beta_C + \delta_C N \left\{ 1'_C \beta_C 1_C \right\} \right] \right] \ln(\pi_C) =$$

$$- \left[\begin{matrix} \alpha_C - \delta_C N + \beta_{CO} \ln(\overline{P}_O) + \delta_C N \beta'_O \ln(\overline{P}_O) + \\ \left[\gamma \ln(N^*) + \Gamma A + (\delta_O N)' \ln(\overline{P}_O) \right] \left[\beta_C + \delta_C N \left\{ 1'_C \beta_C 1_C \right\} \right] \end{matrix} \right] - \varepsilon_C$$

were the "C" and "O" subscripts identify the censored and observed portions of each vector, respectively. We can simplify (9) by modifying the censored good's error terms where we let

$\beta^* \equiv \beta_{CC} + \delta_C N \beta'_C + (\delta_C N)' \left[\beta_C + \delta_C N \left\{ 1'_C \beta_C 1_C \right\} \right]$ and $\eta \equiv \varepsilon_C (\beta^*)^{-1}$. We can then solve for

the (C x 1) vector of reservation prices:

$$(10) \quad \ln(\pi_C) = - \left[\begin{matrix} \alpha_C - \delta_C N + \beta_{CO} \ln(\overline{P}_O) + \delta_C N \beta'_O \ln(\overline{P}_O) + \\ \left[\gamma \ln(N^*) + \Gamma A + (\delta_O N)' \ln(\overline{P}_O) \right] \left[\beta_C + \delta_C N \left\{ 1'_C \beta_C 1_C \right\} \right] \end{matrix} \right] \beta^{-1} - \eta$$

Following Lee and Pitt (1986), the regime switching condition can be represented as:

$$(11) \quad \eta \geq - \ln(\pi_C) - \left[\begin{matrix} \alpha_C - \delta_C N + \beta_{CO} \ln(\overline{P}_O) + \delta_C N \beta'_O \ln(\overline{P}_O) + \\ \left[\gamma \ln(N^*) + \Gamma A + (\delta_O N)' \ln(\overline{P}_O) \right] \left[\beta_C + \delta_C N \left\{ 1'_C \beta_C 1_C \right\} \right] \end{matrix} \right] \beta^{-1} \equiv \Psi^*$$

The likelihood function for this purchase regime can be represented by the product of the conditional density of the original error terms for the purchased goods, conditioned on nonpurchased goods, $g(\varepsilon_{m+1}, \ldots, \varepsilon_K \mid \eta_1 \ldots \eta_m)$, the probability mass of the modified error terms of the zero purchases, $f(\eta_1 \ldots \eta_m)$, and Jacobian transformation from $(\varepsilon_{m+1}, \ldots, \varepsilon_K)$ to

$(x_{m+1}, \ldots, x_K)$ which is a function of the vector of goods and the transformed error terms, $J(x, \eta_1 \ldots \eta_m)$:

$$(12) \quad \int_{\Psi_1^*}^{\infty} \int_{\Psi_2^*}^{\infty} \ldots \int_{\Psi_m^*}^{\infty} J(x, \eta_1 \ldots \eta_m) g(\varepsilon_{m+1}, \ldots \varepsilon_K \mid \eta_1 \ldots \eta_m) f(\eta_1 \ldots \eta_m) d\eta_1 d\eta_2 \ldots d\eta_m$$

Let $I_n(R_c)$ be a dichotomous indicator which equals 1 if the observed consumption of the nth household is associated with purchase regime R_c, zero otherwise. With $l_{n,R_C}(x \mid \theta)$ representing the likelihood function of the nth household and R_c demand regime, the likelihood function (L) value for the N household sample can be represented by:

$$(13) \quad L = \prod_{r=1}^{N} \prod_{R_c} \left[l_{r,R_C}(x \mid \theta) \right]^{I_r(R_c)}$$

Description of Mexican Household Purchase Data

Data for Mexico were obtained from the 1998 Encuesta Nacional de Ingreso y Gastos del Hogar (ENIGH) collected between Aug.-Nov. 1998. This is a nation-wide survey encompassing Mexico's 32 states. Surveyed households maintained weekly diaries of expenditures on a detailed set of food and non-food items. Purchase information includes a disaggregated set of food categories. Household members record their food purchases according to disaggregated set of food categories including not only expenditures but also quantity purchased. In addition, a detailed set of household and household member characteristics are collected.

To avoid problems with respect to the valuation of home produced goods, we limited our current analysis to households that resided in towns with a population greater than 15,000 persons. We also excluded households that did not record any meat expenditures for at-home consumption during the survey week. As noted in the above discussion of equivalence scales, we need to identify a base household type. Similar to Phipps (1998), we limit our analysis to households with both a male and female head present and where at least one of these heads is between the age of 18 and 65. Given the above, our final sample size was 3,610 households. The data described in Table 2 was obtained from this data set.

Table 3 provides an overview of household size and composition of our sample households. Mean household size, given we limit our analysis to two-parent households, was 4.5 with an average 1.9 children under the age of 18. The extended nature of Mexican households is evidenced by the fact that 38% of the sample households have at least one other adult present in the household. Approximately 25% of the households had 7 or more members. Table 3 also provides a summary of the demographic characteristics used in the scaling function. These included the ownership of a refrigerator/freezer (REFRIG) and 8 regional dummy variables.

Examining the Structure of Meat Demand in Mexico

The general likelihood function represented by (12) and (13) was applied to our Mexican household data in an analysis of the consumption of the 5 meat/fish commodities detailed in Table 2. Estimates of the parameters that maximize the likelihood function were obtained by using the GAUSS software system and the BHHH optimization procedure. Given the use of this procedure we use the inverse of the sums of squares and cross-products of parameter gradients across observations evaluated at optimal parameter values as an estimator of the asymptotic covariance matrix of these parameters (Judge, et. al., p.526). Given the complexity of the likelihood function, numerical gradients were used in the estimation procedure. A revised Gibbs sampling technique was used to simulate the truncated multivariate normal distribution associated with the sample's likelihood function represented by (12) and (13).[11] In the estimation of the above model we guarantee that the elements of the equation error term covariance matrix, Σ, will be positive-definite by instead of directly estimating its elements, indirectly estimating these elements by use of the lower triangular matrix, A, where $\Sigma = AA'$, the i,j^{th} of A is a_{ij}, $a_{ij}=0$ for $i<j$ and the a_{ij}'s are estimated parameters.

Table 4 presents estimated parameters when applied to our sample of Mexican households. All of the estimated price and intercept coefficients were found to be individually statistically different from zero. Of particular interest were the scaling function parameters (i.e., the $\gamma_s's$, $\Gamma_1's$, and $\delta_{is}'s$). We find that ownership of refrigerated storage has a significant equivalence scale impact. There is some evidence of significant regional differences in equivalence scale. A likelihood ratio test generated a χ^2-statistic of 102.8 when we estimated a restricted version of the model where all of the regional dummy variables were omitted. This value results in a rejection of the null hypothesis of no regional differences in scaling function values. The "direct" household composition coefficients $(\gamma_s's)$ were both statistically significant. Five of the 10 composition/price coefficients $(\delta_{is}'s)$ were statistically significant.

From the estimated coefficients and similar to Wales and Woodland (1983), Table 5 provides a comparison of compensated own and cross-price elasticities under a number of purchase regimes. In contrast to traditional elasticity derivations, these elasticities incorporate not only the direct impact of a change in price on meat demand but also the indirect effect due to

the scaling function shown in (5). Golan, Perloff and Shen (GPS, 2001) using an earlier (1992) version of this data set use maximum entropy techniques to estimate the parameters of a censored AIDS demand system using the same commodity definitions as used here. They use the usual formulas for evaluating elasticity responses and do not differentiate between intensive and extensive responses to price change. In the top of Table 5 we show estimated compensated demand elasticities evaluated for those surveyed households that purchase all the commodities analyzed here (e.g., intensive responses). The GPS own and cross-price elasticities tend to be higher than those obtained here. This is especially the case for Fish/Shellfish. Golan, Perloff and Shen (2001) find a fairly large net complementary relationship between processed meat with respect to a change in fish price. In contrast we find evidence of a substitute relationship between these goods.

Table 6 reports calculated equivalence scales for two-parent families with alternative combinations of children and other adults present. For illustrative purposes, we set the regional variable, DF, to 1 and assumed the household owned a refrigerator or freezer. Commodity prices were set at their mean value. When interpreting these measures remember our reference household is a childless couple. Thus from the results shown in Table 6, the value of 1.19 for a 3 person household with one child under the age of 18 implies that in terms of the demand for meat, to generate the same level of utility from the consumption of meat as the childless couple, this 3 person household will need to spend 1.19 times the amount spent by the reference household. Moving down the *Child Impact* column we find some evidence of economies although the *differences* in these scale-change values are not significant.

Table 6 can also be used to examine the impact of having other adults in the household on meat demand. Not surprisingly, we find the change in scale values, with the addition of adults to the household, are greater than that of children. For example, when adding another adult to a 2-person household there is a change in scale function of 0.29. This compares to a 0.19 change when adding a child to a 2-person household. The 0.29 value is much less than the expected 0.5 value (given the definition of our base household) indicating some economies when additional adults are added. Notice, that the adult induced scale change increases with the number of children and approaches 0.50 for having an additional adult in a household with 5 children present.

As a comparison, in an analysis of the impact of children on the costs of food, clothing, shelter and transportation by Canadian households, Phipps(1998) found that compared to a childless couple, the addition of one child to the household resulted in a relative equivalence scale value of children of 1.16 for one child, 1.28 for two and 1.38 for three. Phipps and Garner (1994) estimate food equivalence scales for Canada and the U.S. using a series of Engel curves. Unfortunately, they examine the impact of household size on food expenditures regardless of whether these additional members are adults or children. Using a two-person household as a base, they obtain relative food equivalence values of 1.33 and 1.68 for 3 and 4-person households in the U.S. and 1.36 and 1.73 for Canadian households, respectively (p.10-11). Similarly, Blaylock (1991) presents food equivalence values for different size households regardless of age of additional members. Using a 2-person household as a base, he obtains relative equivalence measures of 1.22 and 1.51 for 3 and 4-person households, respectively. From our analysis we obtain scaling function values of 1.19 and 1.35 for 1 and 2-child household respectively. For a 3 and 4-adult household we obtain relative scale values of 1.29 and 1.54, respectively. For a 4-person household with 1 other adult and 1 child we estimate a relative scale value of 1.54 .

Conclusions

The present analysis represents a first step in the analysis of Mexican food demand that is based on disaggregated commodity definitions. The current application allows us to quantify the differential impacts on food demand of household members that differ by age. Our analysis is based on a theoretically consistent method for deriving endogenously determined equivalence scale measures. For this analysis we examine the demand for meat and fish for at-home consumption. An important result we find for potential U.S. exporters is that for the commodities analyzed, conditional own price elasticities were consistently found to be inelastic regardless of purchase regime. We also find evidence that household "size" as represented by the number of scale equivalents indicates that households adjust purchasing behavior such that they realize scale economies for larger number of adult equivalents. This may have implications of potential retailers who may want to offer product at reduced cost/unit so as to attract larger "size" households.

To complete this research we need to: (i) add additional age categories to the currently used "children" and "other adult" member categories, (ii) differentiate household members by gender (iii) incorporate these changes within a complete food demand system (e.g., not just meat demand) and (iv) recognize the importance of food-away-from home for urban residents. Early attempts at adding more detail to the age and addressing the gender dimensions have proven unsuccessful. This may be due to the low number of households with particular age/gender compositions as the number of age categories increase. Alternative functional forms for the equivalence scale speciation will be attempted. Given the limited size of our demand system, extension to a large system (e.g., 10 equations) could prove to be a problem given that the number of parameters to estimate increases in a nonlinear manner with the number of food groups. Initial attempts at estimating a 10-equation demand system (without endogenous scales) have been successful given our use of simulated maximum likelihood techniques although it takes many days of computation using a relatively small sample to obtain parameter estimates. If we are able to overcome the above problems, we will have a method for which we can examine the relative welfare of households of differing compositions when there welfare is defined on their food purchases.

Table 1. Importance of Mexico as a Destination for U.S. Agricultural Exports (1990-2000)

	1990	1991	1992	1993	1994	1995	1996	1997	1998	1999	2000
World U.S. Ag. Exports ($Mil)	39,517	39,392	43,132	42,911	46,244	56,348	60,445	57,245	51,829	48,485	51,580
U.S. Ag. Exports to Mexico ($Mil)	2,560	3,008	3,802	3,619	4,593	3,540	5,447	5,184	6,163	5,634	6,545
% to Mexico	6.5	7.6	8.8	8.4	9.9	6.3	9.0	9.1	11.9	11.6	12.7
Distribution of U.S. Agricultural Exports to Mexico (%)											
Animal & Animal Prod.	26.0	37.4	33.1	32.6	29.7	23.3	20.0	29.7	27.2	27.9	28.7
Grains & Feeds	37.5	24.6	27.9	24.5	26.7	30.0	38.0	22.5	26.6	28.0	26.1
Fruits and Prep.	1.8	1.9	2.0	3.1	4.0	2.4	1.7	2.3	2.1	3.4	3.8
Fruit Juices	0.1	0.2	0.2	0.2	0.3	0.2	0.1	0.2	0.2	0.3	0.5
Nuts & Prep.	0.7	0.9	1.0	1.0	1.0	0.9	0.8	0.8	0.8	1.1	1.2
Veg. & Prep.	7.2	4.0	4.2	4.8	5.4	4.0	4.6	5.4	7.0	6.7	7.1
Oilseeds & Prod.	12.7	17.5	18.8	18.1	18.5	23.5	20.2	23.0	18.7	18.7	15.8
Other Ag. Prod.	14.1	13.6	12.8	15.8	14.4	15.7	14.6	16.2	17.4	14.0	16.9

Source: Various issues of FATUS.

Table 2. Overview of Weekly Mexican Per Capita Food Purchases

Commodity	Mean Per Capita Expenditure (Peso)	% of Total Expenditures	% Households Purchasing	Mean Per Capita Expenditure by Purchasing Households (Peso)	Std. Dev. of Conditional Expenditures (Peso)
Total Food At Home	54.6	100.0	100.0	54.6	32.1
Beans	2.0	3.6	53.9	3.7	11.2
Cheese	2.0	3.6	47.0	4.1	3.7
Fruits	3.2	5.8	14.5	4.9	5.1
Grains	10.7	19.5	98.2	10.8	7.0
Fluid Milk	6.7	12.2	81.7	8.2	6.7
Non-Alcoholic Bev.	5.5	10.1	77.7	7.1	5.9
Vegetables	6.1	11.1	89.8	6.7	5.1
Meat/Fish Expenditures					
Beef	7.7	14.1	71.9	10.7	7.9
Pork	2.0	3.6	28.6	6.9	5.0
Poultry	4.5	8.2	62.8	7.1	4.9
Processed Meat	3.1	5.7	57.8	5.3	4.8
Fish/Shellfish	1.5	2.7	22.9	6.4	6.6
Total Meat/Fish	18.7	34.2	93.0	20.1	14.6

Source: 1998 ENIGH, Urban Households, Male/Female Adult Heads Present and have positive food-at-home expenditures.

Table 3. Household Composition and Other Household Characteristics

Frequency Distribution					
Household Size $\mu=4.6$, $\sigma=1.8$		Non-Head Adults $\mu=1.9$, $\sigma=1.5$		Children < 18 Years $\mu=0.7$, $\sigma=1.2$	
Category	%	Category	%	Category	%
2	7.2	0	61.7	0	17.1
3	17.5	1	17.6	1	24.0
4	27.8	2	11.3	2	29.6
5	23.5	3	6.2	3	17.8
6	12.5	4	2.2	4	7.1
7	6.0	5	0.8	5	2.6
>7	5.5	>5	0.3	>5	1.8

Variable	Description	Mean
REFRIG	Household Owns Rrefrig../Freezer (0/1)	83.6
Regional Dummy Variables (State of Residence)		
DF*	Distrito Federal, Estado de Mexico and Metropolitan Areas around Mexico City	32.4
NW	Baja California, Baja California Sur, Sonora and Sinaloa	8.3
NE	Coahuila, Chihuahua, Nuevo Leon and Tamaulipas	11.1
NC	Durango, San Luis Potosi, Queretaro and Zacatecas	5.6
WEST	Nayarit, Jalisco, Colima, Guanajuato and Michocacan	19.5
CENTRAL	Aguascalientes, Hidalgo, Morelos, Puebla and Tlaxcala	8.8
SOUTH	Guerrero, Oaxaca and Veracruz	6.1
SE	Yucatan, Tabasco, Quintana Roo, Chiapas and Campeche	8.3

Note: *The region DF used as the base region. μ represents sample mean and σ sample standard deviation.

Table 4. Simulated Maximum Likelihood Parameter Estimates

Commodity	Coefficient	Std. Dev.	Commodity	Coefficient	Std. Dev.
Intercept					
Beef	-0.9217[a]	0.0296	Processed Meats	0.1307[a]	0.0144
Pork	0.2542[a]	0.0116	Seafood/Fish	0.3182[a]	0.0099
Poultry	-0.7814[a]	0.0284			
Price Coefficients					
Beef/Beef	0.4516[a]	0.0132	PrcMt/Poultry	-0.0270[a]	0.0034
Pork/Beef	-0.0584[a]	0.0028	PrcMt/PrcMt	0.1337[a]	0.0033
Pork/Pork	0.1669[a]	0.0045	Seafood/Beef	-0.0282[a]	0.0016
Poultry/Beef	0.0139[a]	0.0052	Seafood/Pork	-0.0151[a]	0.0006
Poultry/Pork	-0.0472[a]	0.0024	Seafood/Poultry	-0.0222[a]	0.0014
Poultry/Poultry	0.3844[a]	0.0114	Seafood/PrcMt	-0.0154[a]	0.0007
PrcMt/Beef	-0.0326[a]	0.0034	Seafood/Seafood	0.0817[a]	0.0024
PrcMt/Pork	-0.0275[a]	0.0012			
Demographic Variables					
REFRIG	0.5738[a]	0.0399	WEST	-0.0194	0.0350
NW	0.4476[a]	0.0528	CENTRAL	-0.2259[a]	0.0518
NE	0.0646	0.0515	SOUTH	-0.1413[a]	0.0489
NC	0.0689	0.0729	SE	0.2460[a]	0.0427
Age Composition					
NumKids	0.4377[a]	0.0382	NumAdults	0.6143[a]	0.0391
Beef_NumKids	-0.0053	0.0124	Beef_NumAdults	-0.0398[a]	0.0106
Pork_NumKids	0.0087[a]	0.0041	Pork_NumAdults	0.0051	0.0036
Poultry_NumKids	-0.0102	0.0120	Poultry_NumAdults	0.0209[b]	0.0104
PrcMt_NumKids	0.0025	0.0059	PrcMt_NumAdults	0.0128[a]	0.0057
Fish_NumKids	0.0044[b]	0.0023	Fish_NumAdults	0.0010	0.0020

Note: Due to space limitations, the elements of the matrix A, used to derive the error variance covariance matrix, are not presented. These can be obtained from the authors upon request. Given the presence of some very large numbers of household members, in the empirical model we use the inverse of the transformed member count variables in the direct age component of the scaling function. [a] represents statistical significance at the .01 level and [b] at the .05 level.

Table 5. Hicksian Elasticities Under Alternative Purchase Regimes With
A Comparison to Previous Elasticity Estimates

Purchase All Commodities

	Study	Beef	Pork	Poultry	Processed Meat	Fish
Beef	Current	-0.103	0.016	-0.043	0.046	0.084
	GPS	-0.596	0.187	0.228	0.015	0.166
Pork	Current	0.090	-0.268	0.113	0.113	0.179
	GPS	0.550	-0.418	0.081	-0.059	-0.153
Poultry	Current	0.033	-0.019	-0.050	0.007	0.029
	GPS	0.263	0.034	-0.402	0.111	-0.006
Processed Meat	Current	0.128	0.122	-0.050	-0.357	0.157
	GPS	0.041	-0.052	0.255	-0.706	0.462
Fish	Current	0.150	0.236	-0.120	0.182	-0.448
	GPS	1.236	-0.400	-0.034	-1.285	-2.088
Expenditure	Current	0.844	1.345	0.402	1.257	1.730
	GPS	1.305	1.149	0.745	0.542	1.247

Note: GPS identifies the elasticities reported by Goaln, Perloff and Sen (2001).

Purchase Only Meat Products

	Beef	Pork	Poultry	Processed Meat
Beef	-0.184	0.096	0.016	0.072
Pork	0.209	-0.317	-0.046	0.154
Poultry	0.099	0.050	-0.188	0.039
Processed Meat	0.209	0.250	-0.071	-0.388
Expenditure	0.903	1.418	0.542	1.472

(continued)

Table 5. Hicksian Elasticities Under Alternative Purchase Regimes (continued)

Do Not Purchase Red Meat

	Pork	Poultry	Processed Meat	Fish
Pork	-0.310	0.022	0.137	0.152
Poultry	0.090	-0.175	0.039	0.046
Processed Meat	0.262	-0.005	-0.380	0.124
Fish	0.364	-0.052	0.152	-0.464
Expenditure	1.243	0.541	1.162	1.472

Do Not Purchase Processed Meats

	Beef	Pork	Poultry	Seafood
Beef	-0.118	0.047	-0.058	0.129
Pork	0.112	-0.282	-0.109	0.279
Poultry	0.051	0.024	-0.146	0.071
Fish	0.194	0.283	-0.030	-0.447
Expenditure	0.862	1.447	0.482	1.568

Do Not Purchase Red Meat or Seafood

	Pork	Poultry	Processed Meat
Pork	-0.254	0.081	0.173
Poultry	0.179	-0.255	0.077
Processed Meat	0.386	-0.019	-0.370
Expenditure	1.307	0.649	1.351

Do Not Purchase Seafood or Processed Meat

	Beef	Pork	Poultry
Beef	-0.232	0.165	0.067
Pork	0.279	-0.330	0.051
Poultry	0.171	0.149	-0.320
Expenditure	0.957	1.404	0.776

Table 6. Comparison of Simulated Relative Adult Equivalence Scales for Alternative Household Compositions

No. of Kids	Number of Other Adults in the Household											
	0			1					2				
		Child Impact			Child Impact		Adult Impact			Child Impact		Adult Impact	
	Relative to Base	Scale Change	T-Value	Relative to Base	Scale Change	T-Value	Scale Change	T-Value	Relative to Base	Scale Change	T-Value	Scale Change	T-Value
0	1.00	-----	-----	1.29	-----	-----	0.29	14.32	1.54	-----	-----	0.25	6.12
1	1.19	0.19	10.42	1.54	0.25	6.42	0.35	8.90	1.84	0.30	4.65	0.30	4.72
2	1.35	0.16	4.44	1.74	0.20	3.40	0.39	6.47	2.08	0.24	2.74	0.34	3.82
3	1.49	0.14	2.65	1.92	0.18	2.18	0.43	5.27	2.30	0.22	1.85	0.38	3.28
4	1.61	0.12	1.82	2.08	0.16	1.56	0.47	4.57	2.49	0.19	1.36	0.41	2.92
5	1.73	0.12	1.36	2.22	0.14	1.19	0.50	4.10	2.66	0.17	1.06	0.44	2.67

Note: The *Relative to Base* column is the ratio of the simulated scaling function value to the simulated for a household with a male and female head present and where there are no children or other adults present. For example, the value of 1.29 with the addition of 1 other adult indicates a 29% increase in scaling function value over the base household. The *Child Impact* columns are calculated as the difference between relative scaling function values with changes in number of children. For example, for households with 1 additional adult and 2 children, the child impact value of 0.20(1.74-1.54). The *Adult Impact* columns are calculated as the difference between relative scaling function values for the same number of children but larger number of other adults. For example, the relative adult impact for households with 1 additional adult and 2 children is calculated as 0.39=(1.74-1.35). All prices are evaluated are the sample mean values. The t-values are derived from the approximate standard errors of the relative scale values and are used to test whether the current impact value is statistically different from the previous value where the previous value will depend on whether one is concerned with the child or adult impacts as described above. For this table we assume the household resides in the DF region and owns a refrigerator or freezer.

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Footnotes

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[1] This sample is not representative of Mexican households as we limit our analysis to urban households with a male and female head present.

[2] When applied to household income, adult equivalence scales are employed to adjust household budgets to permit welfare comparisons across households differing in size and composition (Lazear and Michael, 1980). For a review of the methodological issues involved with the estimation of adult equivalence scales for welfare evaluation refer to Blaylock (1991).

[3] For a detailed discussion of adult equivalent scales and demand system estimation, refer to Lewebl (1997)

[4] Similar to Golan, Perloff and Sen (2001), we assume that meat and other goods are separable in the household's utility function. Alson and Chalfant (1987) obtained mixed results in terms of whether seperability holds in Australian meat purchases. In the analysis of U.S. meat purchases, Moschini, Moro and Green (1994) find evidence of seperability.

[5] Refer to Gould, Cox and Perali (1991) for an example of endogenously estimated scaling functions within a demand systems framework.

[6] The processed meat category include such foods as ham, bacon, wieners, chorizo, other smoked/seasoned meats, dried beef/jerky, salami, bologna, etc.

[7] The assumption of equivalence scale exactness implies that this measure is only a function of the demographic characteristics and prices and is independent of utility level.

[8] This form is used so as to allow for the use of logarithms even with zero valued member count variables.

[9] In matrix notation this equation can be represented as:

$$(6')\quad W = \frac{1}{D}\left[\alpha - \delta N + \beta \ln\left(\overline{P}\right) + \delta N\left(\left(1_K \beta\right)\ln\left(\overline{P}\right)\right) + \ln d\left(A,\overline{P}\right)\left(\beta 1_K' + \delta N\left(1_K \beta 1_K'\right)\right) + \varepsilon\right]$$

where $1_K = [1,\ldots.1]$ of dimension K, $\overline{P} = \dfrac{P}{M}$, $\ln d\left(A,\overline{P}\right) = \gamma \ln\left(N*\right) + \Gamma A + \left(\delta N\right)'\ln\left(\overline{P}\right)$, and

$$D = -1 + \left(1_K \beta\right)\ln\left(\overline{P}\right) + \ln d\left(A,\overline{P}\right)\left(1_K \beta 1_K'\right).$$

[10] For a more detailed derivation of these shadow prices, refer to Lee and Pitt (1986).

[11] For the simulated maximum likelihood procedure we used the FGIBBS procedure developed for the GAUSS software system by Dr. Ron Mittlehammer and Dr. Maher Hasan at Washington State University. The FGIBBS procedure is used to create a matrix of (pseudo) random variables distributed truncated multivariate normal using an improved Gibbs sampler. The main idea behind the new sampler is to utilize the fact that we usually use the Gibbs sampler with problems where one continuously draws from a truncated normal until one gets convergence. Instead of throwing a way the old draws we keep the unbiased ones of them according to the new mean and bounds of the conditional normal by utilizing fast and accurate procedure to check for the unbiased draws. Since in most cases, after few iterations, the change in the parameters will become very small it is expected that we will keep many of these draws and the need for new draws obtained from the Gibbs sampler will be minimal. For a review of this procedure refer to Hasan and Mittlehammer (2002) and to Hasan (2001).