RURAL ECONOMY

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Potato Consumption in Canada: Is it Becoming a Normal Good?

The notion that potato is an inferior good dates back to the mid-nineteenth century when British economist Robert Giffen (and later Paul Samuelson) asserted that potato constituted a Giffen good in historic Ireland (McDonough and Eisenhauer). Potato has become a leading practical example of an inferior good, if not a Giffen good (Rosen). By definition, consumption of potato would decrease with income, if it is an inferior good. While potato consumption has been found to be negatively related to GNP per capita in the OECD countries (Andersson and Senauer), recent consumption trends in Canada and the United States seem to suggest the opposite. Per capital potato consumption in Canada fluctuated between 60 kg and 80 kg during 1978-97, but stabilized around 77 kg after 1993 (figure 1). During the same period, consumption in the United States increased steadily between 1978 and 1997, rising from 54.3 kg in 1978 to 64.5 kg per capita in 1997 (figure 1). These consumption trends raise an interesting question: is potato still an inferior good in North America? The answer to this question is important because potato remains an important food item in many countries and many domestic agricultural as well as international trade policies are centered on potato. Such policies would be misdirected if they were based on the misbelief that potato is an inferior good, without a rigorous and robust empirical support for that belief. The objective of this paper is to estimate and evaluate the demand elasticities for potato products in Canada. The survey data we use allow investigation of the demand for fresh as well as other forms of potato products.

As Huang and Bouis point out, an analysis of the simple correlation between aggregate consumption and per capita income does not necessarily reveal the true relationship between consumption and its contributing factors as other conditioning variables such as prices and demographic characteristics are not included. While the effects of prices and income can be masked by collinearity among these variables in time series, more accurate estimates of price and income effects can be obtained from cross-sectional data. To estimate the demand elasticities for potatoes, we use data from the 1996 Family Food Expenditure Survey (FFES) collected by Statistics Canada.

Fresh potatoes have lost market share in recent years not only to processed potato products but also to other staple foods such as rice, pasta and bread (Richards, Kagan and Gao). To capture the impacts of potential substitution among potato products, we use a demand system framework to investigate the demand for three forms of potato products: fresh, frozen, and dried/chipped potatoes. In addition, since rice, pasta and bread are potential substitutes among the Canadian staples, these products are also included in the system. Cereals and other grain products were often included in previous analysis of demand for potatoes (Gao, Wailes and Cramer; Richards, Kagan and Gao).

The use of household-level survey data is often hindered by the occurrence of zero expenditures in the sample. Such zero observation or 'limited dependent variable' issues arise as households participating in the survey typically do not report consumption of all food products during the survey period. A number of censored demand system estimators have been proposed in the literature. The maximum-likelihood procedures of Lee and Pitt (1986, 1987) and Wales and Woodland involve evaluations of multiple probability integrals and for that reason applications of these procedures have remained scanty. A large body of demand studies was based on the two-step procedure of Heien and Wessells, and two additional two-step procedures have been proposed in the literature (Perali and Chavas; Shonkwiler and Yen). These two-step estimators however are known to be inefficient, relative to maximum-likelihood estimators, and proper statistical inference generally involves correction of the standard errors for the second-step estimates, which can be cumbersome for nonlinear demand systems. To overcome these computational complexities in maximum-likelihood and two-step estimation, we use a parsimonious procedure known as the quasi maximum-likelihood (QML) approach, initiated by Avery and Hotz and Avery, Hansen and Hotz in the multivariate probit literature. In the QML approach, the likelihood function is approximated by combining a sequence of bivariate Tobit likelihoods, thus avoiding the need to evaluate multiple probability integrals while allowing cross-equation error correlation and parametric restrictions.

The next section presents the demand system estimated in the study. This is followed by a description of the econometric procedure, data and estimation results. The last section concludes.

The Translog Demand System

In this study we investigate demand for potato and grain products, which are assumed to be weakly separable from all other goods in the consumption bundle. The demand system is derived from the translog utility function (Christensen, Jorgensen and Lau)

$$\log V(v;\theta) = \alpha_0 + \sum_{i=1}^{n} \alpha_i \log v_i + \frac{1}{2} \sum_{i=1}^{n} \sum_{j=1}^{n} \beta_{ij} \log v_i \, v_j \,, \tag{1}$$

where $\log V(v;\theta)$ is indirect utility, $v = [v_1, v_2, \cdots, v_n]'$ is a vector of income-normalized prices, and θ is a vector containing all parameters, namely α_0, α_i and β_{ij} $(i, j = 1, 2, \cdots, n)$. Applying *Roy's identity* to (1) gives the translog demand system

$$w_i = \frac{\alpha_i + \sum_{j=1}^{n} \beta_{ij} \log v_j}{\sum_{j=1}^{n} \alpha_j + \sum_{k=1}^{n} \sum_{j=1}^{n} \beta_{kj} \log v_j}, \;\; i = 1, 2, \cdots, n, \tag{2}$$

where $w_i = v_i q_i$ is expenditure share and q_i is quantity for good *i*. Homogeneity is implicit in equations (1) and (2) by use of normalized prices v. Demographic variables are incorporated in (2) by letting

$$\alpha_i = \sum_k \alpha_{ik} z_k \,, \tag{3}$$

where z_1 is unity. Parametric restrictions imposed include symmetry $(\beta_{ij} = \beta_{ji} \; \forall \; i, j)$ and 'adding-up':

$$\sum_{i=1}^{n} \alpha_{i1} = -1, \; \sum_{i=1}^{n} \alpha_{ik} = 0 \;\; \forall \; k \geq 2. \tag{4}$$

However, the restrictions (4) guarantee adding-up only in the absence of censoring. The issue of adding up in the type of censored model considered is addressed in the next section.

Estimation of a Censored System

Denote the deterministic component of the demand share equation for good *i* as $f_i(\theta)$. The system of censored demand equations we consider is a nonlinear extension of

Amemiya:

$$w_i = f_i(\theta) + \varepsilon_i \quad \text{if} \quad f_i(\theta) + \varepsilon_i > 0$$
$$= 0 \qquad \text{otherwise} \qquad i = 1, 2, \cdots, n, \tag{5}$$

where ε_i $(i = 1, 2, \cdots, n)$ are error terms. This additive stochastic structure in (5) is consistent with the general error model of McElroy. In the presence of censoring, the right-hand side of the system (5) no longer adds up to unity even if $\sum_{i=1}^{n} \varepsilon_i = 0$. To accommodate the adding-up restriction we estimate the first n – 1 equations in the system and treat the nth equation as a residual demand.[1] Elasticities for the nth good are then calculated using the adding-up restriction. Note that full-information maximum-likelihood (FIML) estimates are not invariant to the equation excluded. To discuss the estimation procedure, consider, without loss of generality, a demand regime with observed (n–1)th vector $w = [0, \cdots, 0, w_{\ell+1}, \cdots, w_{n-1},]'$ in which the first ℓ goods are not consumed. Assume random error vector $e \equiv [\varepsilon_1, \varepsilon_2, \cdots, \varepsilon_\ell \mid \varepsilon_{\ell+1}, \varepsilon_{\ell+2}, \cdots, \varepsilon_{n-1}]' \equiv [e_1', e_2']'$ is distributed as (n–1)-variate normal $e \sim N(0, \Sigma)$, where Σ is a constant and contemporaneous covariance with entries $\sigma_{ij} = \rho_{ij}\sigma_i\sigma_j$ $(i, j = 1, 2, \cdots, n-1)$, σ_i's are error standard deviations and ρ_{ij}s are correlation coefficients. Then, the likelihood contribution of this regime is

$$L_c(w) = g(e_2) \int_{\{e_1 : e_1 \leq u\}} h(e_1 \mid e_2)\, de_1, \tag{6}$$

where $u = -[s_\ell(\theta), \cdots, s_\ell(\theta)]'$, $g(e_2)$ is the marginal probability density function (pdf) of e_2 and $h(e_1 \mid e_2)$ is the conditional pdf of e_1 given e_2. Both $g(e_2)$ and $h(e_1 \mid e_2)$ are normal pdf's, with appropriate moments (means and covariances) following from the normality of e (Kotz, Balakrishnan and Johnson). Thus, the integral in (6) can be evaluated as a ℓ-dimensional normal cdf. The likelihood contribution (6) reduces to one extreme regime of no censoring, with likelihood contribution corresponding to the (n–1)-dimensional pdf of e, namely $f(e)$. The other extreme regime is one in which all (n–1) goods are zeroes, for which the likelihood contribution involves integration of $f(e)$ over the entire (n–1)-vector e. The sample likelihood function is the product of the likelihood contributions (6) over the sample.

In this study we consider a system of six equations and our sample contains over one half of observations with zeroes in four or more commodities, which requires evaluation of four-dimensional normal probability integrals or higher in FIML estimation. To overcome the computational complexity, we use a procedure known as the quasi maximum-likelihood (QML) approach. The QML procedure, initiated in the estimation of multivariate probit (Avery, Hansen and Hotz; Avery and Hotz) and used in subsequent applications of censored linear systems (Harris and Shonkwiler; Yen and Lin), approximates the full-information likelihood function (6) with a sequence of bivariate Tobit likelihoods. We applied the procedure to the censored nonlinear system considered in this study. Denote $z_i = [w_i - f_i(\theta)]/\sigma_i$ and $z_j = [w_j - f_j(\theta)]/\sigma_j$, and define a dichotomous indicator $I(w_i = 0, w_j > 0)$ which equals one if $w_i = 0$ and $w_j > 0$ and zero otherwise, etc., then the bivariate Tobit likelihood for equations i and j for an observation is

$$
\begin{aligned}
L_{ij} = & \left\{\Psi(z_i, z_j, \rho_{ij})\right\}^{I(w_i=0, w_j=0)} \left\{\sigma_i^{-1}\sigma_j^{-1}\psi(z_i, z_j, \rho_{ij})\right\}^{I(w_i>0, w_j>0)} \\
& \times \left\{\sigma_j^{-1}\phi(z_j)\Phi\left[(z_i - \rho_{ij} z_j)/(1-\rho_{ij}^2)^{1/2}\right]\right\}^{I(w_i=0, w_j>0)} \\
& \times \left\{\sigma_i^{-1}\phi(z_i)\Phi\left[(z_j - \rho_{ij} z_i)/(1-\rho_{ij}^2)^{1/2}\right]\right\}^{I(w_i>0, w_j=0)},
\end{aligned} \tag{7}
$$

where $\phi(\cdot)$ and $\Phi(\cdot)$ are univariate standard normal pdf and cumulative distribution function (cdf), respectively, and $\psi(\cdot,\cdot,\cdot)$ and $\Psi(\cdot,\cdot,\cdot)$ are bivariate standard normal pdf and cdf. Appending a subscript to the bivariate likelihood to index observation *t*, the quasi-likelihood function for a sample of *T* observations is

$$
L = \prod_{t=1}^{T} \prod_{i=1}^{n-2} \prod_{j=i+1}^{n-1} L_{ijt}. \tag{8}
$$

Censoring in the dependent variable has to be accommodated when calculating elasticities. This can be accomplished by a procedure parallel to that of McDonald and Moffitt for the linear Tobit model. For each product *i*, the unconditional mean of the dependent variable (expenditure share) is

$$
E(w_i) = \Phi[\, f_i(\theta)/\sigma_i]\, s_i(\theta) + \sigma_i \phi[\, f_i(\theta)/\sigma_i]. \tag{9}
$$

Elasticities can be derived by differentiating (9). Detail elasticity formulas are available from the authors.

Data

Data used in this paper are compiled from the 1996 Family Food Expenditure Survey (FFES) collected by Statistics Canada (Statistics Canada). The number of households selected was 10,695, each of which was interviewed in two consecutive weeks. Although one might contemplate treating replicates of the same household as separate observations, doing so would cause statistical problems. This is because, for households with complete two-week data, the values of most explanatory variables do not vary from one week to the next. Consequently, variations in weekly consumption are likely to be picked up by the error terms, causing (inter-temporal) correlation among the errors that is hard to accommodate with the short time period (two weeks) in the ‘panel’. To avoid such statistical complications, data were aggregated over the two-week period. Such aggregation is helpful, as one week may be too short for revealment of preference. Two-week data should also exhibit less occurrence of zero expenditures caused by infrequency of purchases.

About 973 households containing zero expenditures for all six products were excluded from the sample because expenditure shares are not defined for these households.[2] A small number of households with missing data for selected variables were also excluded. The final sample includes 9,790 observations.

The FFES data contain a detailed list of household food expenditures on numerous household food items. There are four types of potato products: fresh, chips, dried (dehydrated) and frozen potatoes. Approximately 43% of the households reported fresh potato consumption during the two survey weeks, whereas 39% consumed potato chips, 18% consumed frozen potato, and only 2% consumed dried potato. Due to the small number of positive observations (and therefore lack of variation) in dried potato consumption, potato chips and dried potato were aggregated into one category. Also included in the system are three other staple food products: rice, bread and pasta. Approximately 43% of the sample reported consumption of rice during the two-week period, whereas 39% consumed bread, and 18% consumed pasta.

Table 1 presents the frequency distribution of zeros among the six expenditure shares. Only 146 households (about 1.5% of the sample) reported consumption of all six products during the two-week period. With the last equation (pasta) excluded, which is the more appropriate distribution in assessing computational burdens, the distribution shows that 176 households (1.8%) consume all five of the products, and that 3,047 households (31.1% of sample) contain four or more zeros, which would have called for evaluation of four and five-level integrations for FIML estimation.

For each of the six products price is approximated by the unit value, derived as the reported expenditure (in cents) divided by the quantity purchased (in grams). For households which did not purchase during the survey period and therefore for which no price data were available for the product, regional/seasonal average prices are used.[3] To account for heterogeneous preference, a number of socioeconomic and demographic factors are also used. These variables include: household composition in four age categories and race, as well as age, gender, marital status and education of the household head. Table 2 presents the descriptive statistics for the product categories and table 3 presents the definitions and sample statistics of demographic variables.

Estimation Results

QML estimation of the censored translog demand system is carried out by maximizing the quasi likelihood function (8), using the ‘maxlik’ procedure in the Gauss programming language. Analytic gradients of the quasi-likelihood function were used, and numerical optimization is done with the Broyden-Fletcher-Goldfarb-Shanno (BFGS) algorithm (Luenberger).[4] Finally, robust covariance matrix of the QML estimates is calculated using White's heteroscedasticity-consistent procedure.

The estimation results are presented in table 4. In assessing the parameter estimates, 17 (or about 34%) of the 50 demographic parameter estimates (α_{ik}'s) and 13 (61.9%) of the 21 quadratic price coefficients (β_{ij}'s) are significant at the 5% level of significance. All estimated error standard deviations (σ_i's) and all but one of the error correlation coefficients (ρ_{ij}'s) are significant at a significance level of 1% (p-values $<$ 0.0001). Overall, one half of all parameter estimates are significant at the 5% level and 72% are significant at the 10% level. The significance of these demographic variables justifies the accommodation of preference heterogeneity and suggests that household characteristics do play significant roles in determining potato consumption in Canada. Apart from the need to impose cross-equation parametric restrictions, significance of the error correlation coefficients also justifies estimation of the equations in a system vis-à-vis single-equation estimation.

Using the parameter estimates, demand elasticities are calculated by differentiating the unconditional mean (9).[5] Table 5 reports the Marshallian price elasticities as well as expenditure elasticities, along with their standard errors, calculated using the delta method (Ruud, p. 366). All expenditure elasticities are positive and statistically significant at the 5% level of significance. Expenditure elasticities are greater than unity for fresh potato, frozen potato, dried/chipped potato and rice, but are less than unity for bread and pasta. Assuming positive income elasticity for food

(Huang), these positive expenditure elasticities would translate into positive income elasticities.[6] These results would classify the potato products as normal goods, including the fresh potato – an important finding. While fresh potatoes may have played its historic role as a Giffen good, we find no such evidence in the current investigation for Canada. These expenditure elasticities also suggest that, as consumer income grows, the shares of rice and potato products (including fresh potato) would increase faster than those of pasta and bread.

All Marshallian own-price elasticities are significant (at the 5% level), negative and greater than unity (in absolute value). Thus, demands for these products are all price-elastic. These results suggest that potato products in Canada are characterized by downward sloping demand curves and therefore are not Giffen goods. An important marketing implication is that an isolated price decrease in each of these products will increase quantity demanded by a greater proportion, leading to an increase in sales revenue.

Most Marshallian cross-price elasticities are significant at the 5% or 10% level of significance. Fresh potato is a gross substitute to frozen potato, dried/chipped potato, pasta and bread but a gross complement to rice.

Table 6 presents the Hicksian price elasticities. These compensated cross-price elasticities indicate that fresh potato is a net substitute to frozen potato, potato chips, pasta, rice and bread. The different signs in the Marshallian and Hicksian cross-price elasticity between fresh potato and rice indicate that income effect outweighs the substitute effect. On balance, the compensated cross-price elasticities suggest that net substitutability is the more obvious pattern among the six products than net complementarity.

The elasticities with respect to the continuous demographic variables are presented in table 7. The numbers of younger household members (members aged < 15 and 15−24) have positive effects on frozen and dried/chipped potato consumption but negative effects on fresh potatoes and rice. An interesting implication of these elasticities is that lifestyle is an important factors in potato consumption. Households with more middle-aged members (aged 25−64) also consume more dried/chipped potatoes, at the expense of pasta, than others. The effects of age are equally interesting. As a household ages (i.e., headed by an older household head), consumption of dried/chipped and frozen potatoes decreases while consumption of fresh potato and bread increases.

Previous demand estimates appear to vary widely cross studies (Huang; McCracken; Jones and Ward; Guenthner, Levi and Lin; Gao, Wailes and Cramer; Richards, Kagan and Gao). Whereas most of these studies use time series data, Gao, Wailes and Cramer use cross-sectional data from the U.S. Nationwide Food Consumption Survey (NFCS) from 1987-88, a period that is comparable to our study. They reported expenditure elasticities of 0.96, 1.17, 0.88, and 1.04 for rice, potatoes (including frozen, fresh, chips, and dried), bread and pasta, respectively. Our expenditure elasticities for potatoes are comparable to that reported by Gao, Wailes and Cramer, whereas our estimates are much greater for rice and much smaller for pasta. Using the U.S. time series data, Richards, Kagan and Gao also found positive expenditure elasticities for both fresh and frozen potatoes but their magnitudes are much smaller (0.15 and 0.04, respectively) than our estimates. In general, our demand elasticity estimates tend to be higher than those reported in previous studies. These higher elasticities are not surprising because, as more substitutes become available, both the cross- and own-price elasticities of demand are likely to increase. Our results also highlight the importance of using cross-sectional data in estimating the demand elasticities for potatoes.

Conclusion

In the literature, potatoes are typically considered to be an inferior good. This premise would imply that economic growth would bring about a fall in potato demand. This study aims to investigate the roles of prices, income and demographic characteristics in potato consumption, and attempts to determine whether or not potato is still an inferior good in Canada. Analysis is based on the 1996 Family Food Expenditure Survey – the most recent comprehensive household food consumption survey in Canada. To accomplish our objective, a translog demand system is estimated for fresh potato, frozen potato, chipped/dried potato, rice, bread and pasta. The use of household-level data presents an obvious advantage over aggregate time series but it also complicates the econometric methodology. We use a censored demand system estimator to accommodate censoring in the dependent variables. Specifically, to avoid computational (numerical) complexity associated with full-information maximum-likelihood and two-step estimation, we use a procedure known as the quasi maximum-likelihood approach. Most of the price and expenditure elasticities are highly significant. Fresh potato is found to be a normal good, although the income elasticity does not exceed unity, as are the other potato products and staples. Potato products are also found to be price-elastic. These own-price elasticities suggest that price reduction can promote potato sales.

Footnotes

1. The idea of treating the last equation as the residual demand to accommodate adding-up in a censored system is discussed in Pudney (1989).

2. While a sample-selection type of correction might be considered to accommodate selection of these households out of the sample, such sample selectivity would complicate the current framework dramatically. In addition, the small proportion (about 7.4%) of such households would have prevented reliable estimation of a sample-selection equation.

3. Regional/seasonal averages for prices were calculated according to a two-way classification of four seasons and ten provinces in Canada. The literature on missing prices in cross-sectional demand analysis has not settled. While endogenous unit value framework has been considered in the literature, the issues of multiple missing prices in censored systems await theoretical contribution.

4. Analytic gradients are available upon request from the authors.

5. Elasticity formulas are also available from the authors.

6. When the income elasticity for food is around 0.2 (Huang), estimated income elasticities would be 0.32, 0.25, 0.25, 0.19, 0.35 and 0.08 for frozen potato, fresh potato, potato chips, bread, rice and pasta, respectively.

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Figure 1: Per Capita Annual Total Potato Consumption in Canada and the United States, 1978-97

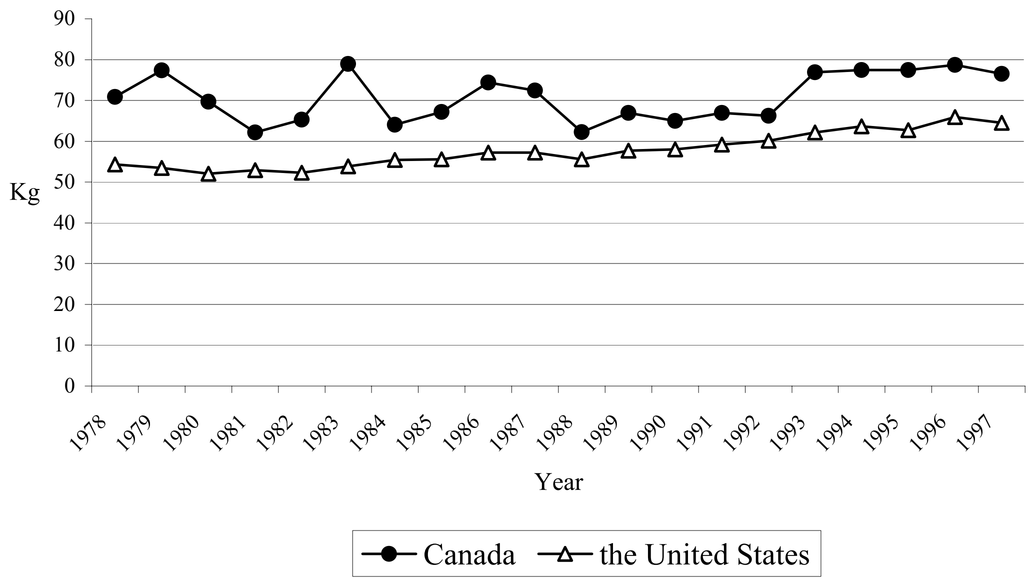

Table 1. Frequencies of Zeros

	All six products		With pasta excluded	
Number of zeros	Frequency	Percent (%)	Frequency	Percent (%)
0	146	1.5	176	1.8
1	689	7.0	924	9.4
2	1619	16.5	2248	23.0
3	2418	24.7	3395	34.7
4	2794	28.5	2891	29.5
5	2124	21.7	156	1.6

Note: Frequency distribution with last equation excluded is more appropriate in evaluating computational burden.

Table 2. Sample Statistics of Product Categories

Variable	Mean	S.D.	% Con-suming
Quantities (kg / two weeks)			
Frozen potato	0.38	1.08	19.04
	(2.02)	(1.67)	
Fresh potato	3.00	6.81	46.04
	(6.51)	(8.82)	
Chips and dried potatoes	0.32	0.66	42.71
	(0.75)	(0.83)	
Bread	3.03	3.54	87.90
	(3.44)	(3.58)	
Rice	0.43	2.21	18.84
	(2.27)	(4.66)	
Pasta	0.87	15.78	48.64
	(1.79)	(1.87)	
Expenditures ($ / two weeks)			
Frozen	0.61	1.59	
	(3.18)	(2.26)	
Fresh	1.73	2.60	
	(4.56)	(3.70)	
Chips/dried	1.95	3.30	
	(6.48)	(5.30)	
Bread	5.70	5.40	
	(5.15)	(6.11)	
Rice	9.70	3.33	
	(4.74)	(4.38)	

Pasta	2.31	3.87
	(4.74)	(4.38)
Shares		
Frozen	0.04	0.11
Fresh	0.14	0.21
Chips/dried	0.14	0.21
Bread	0.48	0.32
Rice	0.05	0.14
Pasta	0.15	0.22
Prices ($ / kg.)		
Frozen	1.81	0.41
Fresh	0.90	0.55
Chips/dried	7.80	2.42
Bread	2.20	0.82
Rice	4.23	1.30
Pasta	3.40	1.64

Note: Rice includes mixes and pasta includes canned pasta products, dried or fresh pasta as well as pasta mixes. Numbers in parentheses are computed from consuming households only.

Table 3. Definitions and Sample Statistics of Demographic Variables

Variable	Definition	Mean	S.D.
Members < 15	Number of members aged < 15	0.53	0.80
Members 15−24	Number of members aged 15−24	0.34	0.64
Members 25−64	Number of members aged 25−64	1.38	0.78
Members ≥ 65	Number of members aged ≥ 65	0.29	0.60
Age	Age of household head	47.33	15.64
Married	Household head is married	0.65	
Female	Household head is female	0.52	
Asian	Household is Asian	0.04	
College	Household head had some college or higher	0.48	

Table 4. Quasi Maximum-Likelihood Estimation of Censored Translog Demand System

Variables	Frozen	Fresh	Chips/dried	Bread	Rice
Demographic variables (α_{ik})					
Constant	1.608**	1.819**	0.280**	–0.438**	2.613**
	(0.451)	(0.684)	(0.121)	(0.113)	(1.094)
Members < 15	–0.080**	0.107**	–0.078*	0.019	0.081**
	(0.033)	(0.052)	(0.046)	(0.018)	(0.041)
Members 15–24	–0.062*	0.090**	–0.116**	0.012	0.063*
	(0.032)	(0.044)	(0.059)	(0.019)	(0.039)
Members 25–64	0.011	–0.025	–0.131*	0.017	0.014
	(0.047)	(0.036)	(0.071)	(0.028)	(0.051)
Members ≥ 65	–0.036	0.008	–0.054	0.053	0.104
	(0.063)	(0.045)	(0.053)	(0.047)	(0.084)
Age	0.072*	–0.111**	0.151**	–0.101**	0.011
	(0.039)	(0.043)	(0.051)	(0.041)	(0.026)
Married	–0.028	–0.088*	–0.037	0.068	–0.096
	(0.056)	(0.050)	(0.048)	(0.043)	(0.077)
Female	0.063	–0.131	–0.011	0.053**	–0.124*
	(0.047)	(0.059)	(0.030)	(0.027)	(0.074)
Asian	0.210*	0.148	0.179*	0.091	–0.548**
	(0.131)	(0.099)	(0.107)	(0.063)	(0.179)
College	0.099*	–0.086*	–0.010	0.022	–0.121*
	(0.059)	(0.047)	(0.030)	(0.023)	(0.074)
Quadratic price terms (β_{ij})					
Frozen	0.284**				
	(0.144)				

Fresh	–0.070	0.248**			
	(0.048)	(0.098)			
Chips/dried	0.089*	–0.024	0.302**		
	(0.051)	(0.025)	(0.122)		
Bread	0.135**	0.072*	0.061**	0.115	
	(0.059)	(0.042)	(0.027)	(0.077)	
Rice	–0.356**	0.001	–0.204**	0.128**	0.435**
	(0.144)	(0.026)	(0.098)	(0.054)	(0.172)
Pasta	0.153**	–0.011	–0.035	–0.390**	0.400**
	(0.049)	(0.025)	(0.029)	(0.113)	(0.165)
Standard dev. (σ_i)	0.375**	0.384**	0.401**	0.342**	0.462**
	(0.010)	(0.005)	(0.006)	(0.003)	(0.012)
Error correlation (ρ_{ij})					
Fresh	–0.162**				
	(0.019)				
Chips/dried	–0.027	–0.220**			
	(0.019)	(0.015)			
Bread	–0.321**	–0.485**	–0.512**		
	(0.015)	(0.010)	(0.009)		
Rice	–0.145**	–0.134**	–0.145**	–0.428**	
	(0.021)	(0.019)	(0.019)	(0.015)	
Log-likelihood	–84972.03				

Note: Asymptotic standard errors in parentheses: ** denotes significance at the 5 per cent level, respectively. Not shown in the table is the parameter estimate for β_{66} (pasta), which is –0.781 with a standard error of 0.222.

Table 5. Marshallian Price Elasticities and Expenditure Elasticities

Products	Frozen	Fresh	Chips	Bread	Rice	Pasta	Expend.
Frozen	–1.548**	0.006	–0.226**	–0.266**	0.355**	0.079	1.599*
	(0.110)	(0.048)	(0.053)	(0.053)	(0.066)	(0.056)	(0.038)
Fresh	0.071**	–1.274**	0.022	–0.080**	–0.010	0.026	1.245**
	(0.033)	(0.034)	(0.023)	(0.025)	(0.027)	(0.028)	(0.019)
Chips/dried	–0.107**	0.015	–1.334**	–0.072**	0.197	0.073**	1.229**
	(0.035)	(0.023)	(0.049)	(0.036)	(0.038)	(0.030)	(0.037)
Bread	–0.019	0.015	0.015	–1.039**	0.033**	0.060**	0.935**
	(0.019)	(0.013)	(0.017)	(0.016)	(0.017)	(0.016)	(0.012)
Rice	0.333**	–0.105**	0.164**	–0.219**	–1.739**	–0.183**	1.749**
	(0.054)	(0.034)	(0.052)	(0.043)	(0.09)	(0.075)	(0.04)
Pasta	0.120**	0.217**	0.227**	0.383**	–0.106**	–1.215**	0.375**
	(0.049)	(0.030)	(0.045)	(0.046)	(0.046)	(0.049)	(0.039)

Note: Asymptotic standard errors in parentheses: ** denotes significance at the 5 per cent level, respectively.

Table 6. Compensated Price Elasticities

Products	Frozen	Fresh	Chips	Bread	Rice	Pasta
Frozen	1.479**	0.235**	–0.007	0.472**	0.439**	0.339**
	(0.110)	(0.049)	(0.052)	(0.048)	(0.067)	(0.056)
Fresh	0.125**	–1.095**	0.192**	0.494**	0.056**	0.228**
	(0.033)	0.036)	(0.024)	(0.023)	(0.027)	(0.028)
Chips/dried	–0.054	0.191**	–1.166**	0.495**	0.261**	0.273**
	(0.035)	0.023)	(0.049)	(0.028)	(0.037)	(0.033)
Bread	0.021	0.150**	0.142**	–0.607**	0.082**	0.212**
	(0.018)	0.012)	(0.017)	(0.016)	(0.017)	(0.015)
Rice	0.409**	0.146**	0.403**	0.587**	–1.647**	0.102
	(0.055)	0.033)	(0.054)	(0.036)	(0.091)	(0.076)
Pasta	0.136**	0.271**	0.278**	0.556**	–0.087*	–1.154**
	(0.049)	(0.029)	(0.047)	(0.037)	(0.046)	(0.051)

Note: Asymptotic standard errors in parentheses: * and ** denote significance at the 5 per cent and 10 percent levels, respectively.

Table 7. Demographic Elasticities

	Members < 15	Members 15–24	Members 25–64	Members ≥ 65	Age
Frozen	0.066**	0.033**	–0.023	0.016	–0.523**
	(0.022)	(0.015)	(0.098)	(0.027)	(0.141)
Fresh	–0.061**	–0.034**	0.038	–0.003	0.566**
	(0.012)	(0.009)	(0.052)	(0.014)	(0.071)
Chips/dried	0.045**	0.043**	0.194**	0.017	–0.764**
	(0.013)	(0.009)	(0.058)	(0.015)	(0.086)
Bread	–0.006	–0.003	–0.015	–0.010	0.305**
	(0.006)	(0.004)	(0.025)	(0.007)	(0.032)
Rice	–0.054**	–0.027*	–0.024	–0.038	–0.063
	(0.021)	(0.014)	(0.089)	(0.025)	(0.169)
Pasta	0.006	–0.001	–0.039*	0.005	0.025
	(0.005)	(0.004)	(0.022)	(0.006)	(0.037)

Note: Asymptotic standard errors in parentheses: * and ** denote significance at the 5 per cent and 10 percent levels, respectively.