

Estimation of a Demand System with Limited Dependent Variables

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Abstract

The study employs the full-information maximum-likelihood method to estimate a censored translog demand system. U.S. household consumption of steak, roast, and ground beef are used to demonstrate the application of the estimation procedure. The proposed methodology produces more efficient estimates than the popular two-step procedures found in demand literature.

Key Words: full-information maximum-likelihood method, two-step procedures, censored translog demand system, demand elasticities.

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Introduction

Red meat consumption in the United States has decreased significantly in the past few decades following a steady downward trend started in the late 1970s. Per capita consumption of beef, accounting for approximately 60% of total red meat, reached an all-time high of 88.8 lbs. in 1976. It dropped about 19% to 72.1 lbs. in 1980 and remained relatively flat in the early 1980s, and then steadily declined from 74.6 lbs. in 1985 to 56.1 lbs. in 1998 (Figure 1). Considerable interest and concern have focused on the trend of declining red meat consumption with special attention given to beef consumption. Smallwood, Haidacher, and Blaylock provided a comprehensive review of the literature on meat demand with broad perspective on significant economic and demographic factors that affect the demand for meat. Chavas (1989) suggested that the changes in meat consumption could be explained mainly by changing meat prices and changing life style of the American consumers.

While numerous studies have focused on the price effects (Chavas, 1983; Dahlgran; Moschini and Meilke; Wohlgenant) based on aggregate time-series data, analysis of the effects of changing life styles, tastes and preferences on meat demand that requires the use of cross-sectional microdata has received little attention. Manchester suggests that demand analyses based on the aggregate time-series data are unsatisfactory because aggregate data usually mask many changes in the groups that make up the whole. Furthermore, analyses based on aggregate economic measures provide price and income elasticities but not shifters for the demand function related to changes in socioeconomic and demographic characteristics of the population.

With the increasing availability of microdata, more recent studies have focused on the effects of demographic characteristics (Capps and Havlicek; Gao and Spreen; Heien and Pompelli; Nayga) and taste change (Gao, Wailes, and Cramer) on demand for disaggregated meat products. Analytical demand studies based on microdata provide better insights on how different groups within the population behave and how the changing relative importance of those groups would have affected the whole economy. Taking the individual household on a micro level, microeconomic models enable provision of better estimation of demand parameters and improvement of forecasts over those that assume average effects for all members of the population based on aggregate data (Manchester). The ability to provide more accurate and improved forecasting of future demands is particularly important to decision makers in the beef industry as well as government officials in formulating sound marketing strategies and public policies.

The analysis of microdata, however, often encountered a limited or censored dependent variable problem. Earlier studies did not address the issues of censored dependent variables (Capps and Havlicek; Heien and Pompelli). It is well known that estimation procedures not accounting for the censored dependent variables produce biased and inconsistent parameter estimates. More recent studies (Gao and Spreen; Gao, Wailes, and Cramer; Nayga) addressed this censoring issue within the framework of demand systems. These studies have typically applied a two-step estimation procedure developed by Heien and Wessells for demand systems with censored dependent variables. The procedure involves a set of probit equations in the first step, and a system of equations augmented by additional regressors (“inverse Mills’ ratios” calculated from the probit estimates) in the second step to account for selectivity biases that may

be present in the demand system. The second-step system of demand equations is estimated by the seemingly unrelated regression procedure.

Two alternative two-step estimation procedures have been proposed recently. Shonkwiler and Yen introduce a procedure based on the unconditional means of the censored dependent variables, which also involves probit in the first step and seemingly unrelated regression in the second step. Using Monte Carlo simulations, Shonkwiler and Yen demonstrate that the alternative two-step procedure performs better than the existing two-step procedure. The procedure of Perali and Chavas involves estimating each equation in unrestricted form using the jackknife technique and then recovering the theoretically restricted demand parameters by minimum distance estimation. We are not aware of any empirical application (except Perali and Chavas) of these recent estimation procedures. As in other two-step estimation procedures, however, these existing two-step procedures generally produce inefficient parameter estimates relative to the full-information maximum-likelihood (ML) estimator.

The main objective of this study is to demonstrate the estimation of a system of demand equations for disaggregated beef products based on a sample of cross-sectional household data collected from the U.S. Department of Agriculture's (USDA) 1987-88 Nationwide Household Food Consumption Survey (NFCS). Specifically, the full-information ML estimation procedure applicable to a censored system of equations will be developed. This methodology will then be applied to estimate U.S. household consumption of three major beef products – steak, roast, and ground beef – based on a translog demand system specification. Given the structural parameter estimates, conditional and unconditional demand elasticities with respect to prices and income will be computed, and policy implications of the empirical estimates will be discussed.

Maximum Likelihood Estimation of a Censored Demand System

Denote the deterministic component of the demand function for good i and observation t as

$f_{it}(x_t, \theta)$, where x_t is a vector of exogenous variables and θ is a vector of parameters. Consider

the system of censored (Tobit) equations (Amemiya), such as

$$(1) \quad \begin{aligned} w_{it} &= f_{it}(x_t, \theta) + \varepsilon_{it} & \text{if } f_{it}(x_t, \theta) + \varepsilon_{it} > 0 \\ &= 0 & \text{otherwise} \end{aligned} \quad i = 1, 2, \dots, n; t = 1, 2, \dots, T,$$

where w_{it} is expenditure share and the error terms $\varepsilon_{1t}, \varepsilon_{2t}, \dots, \varepsilon_{nt}$ are distributed as multivariate

normal with zero mean and constant covariance matrix. As the adding-up restriction of the

demand system implies singularity of the error covariance matrix, estimation must be based on n

$- 1$ equations. Without loss of generality, consider the first $n - 1$ equations for which the first k

goods are zeros. Denote the multivariate normal density of $\varepsilon_t = [\varepsilon_{1t}, \dots, \varepsilon_{kt}, \varepsilon_{k+1,t}, \dots, \varepsilon_{n-1,t}]'$ as

$g(\varepsilon_{1t}, \dots, \varepsilon_{kt}, \varepsilon_{k+1,t}, \dots, \varepsilon_{n-1,t}; \Omega)$, where the covariance matrix

$$\Omega = \begin{bmatrix} \sigma_1^2 & \rho\sigma_1\sigma_2 & \cdots & \rho\sigma_1\sigma_{n-1} \\ & \sigma_2^2 & \cdots & \rho\sigma_2\sigma_{n-1} \\ & & \ddots & \vdots \\ & & & \sigma_{n-1}^2 \end{bmatrix}$$

is positive definite. Then, the contribution to likelihood function of this observation is

$$(2) \quad \int_{-\infty}^{-f_{kt}(x_t, \theta)} \cdots \int_{-\infty}^{-f_{1t}(x_t, \theta)} g[\xi_1, \xi_2, \dots, \xi_k, w_{k+1,t} - f_{k+1,t}(x_t, \theta), \dots, w_{n-1,t} - f_{n-1,t}(x_t, \theta); \Omega] d\xi_1 d\xi_2 \cdots d\xi_k.$$

ML estimation involves evaluation of the multiple integrals in equation (2), which can be

prohibitively expensive in a large demand system with many zeros. In the current application we

consider a system of three beef products, for which the probability integral (2) is greatly

simplified. The sample likelihood function for the three-good case, with the third equation

deleted for estimation, is written as

$$\begin{aligned}
 (3) \quad L = & \prod_{w_{1t}=0, w_{2t}=0} \int_{-\infty}^{-f_{2t}(x_t, \theta)} \int_{-\infty}^{-f_{1t}(x_t, \theta)} g(\xi_1, \xi_2; \Omega) d\xi_1 d\xi_2 \\
 & \times \prod_{w_{1t}=0, w_{2t}>0} \int_{-\infty}^{-f_{1t}(x_t, \theta)} g[\xi_1, w_{2t} - f_{2t}(x_t, \theta); \Omega] d\xi_1 \\
 & \times \prod_{w_{1t}>0, w_{2t}=0} \int_{-\infty}^{-f_{2t}(x_t, \theta)} g[w_{1t} - f_{1t}(x_t, \theta), \xi_2; \Omega] d\xi_2 \\
 & \times \prod_{w_{1t}>0, w_{2t}>0} g[w_{1t} - f_{1t}(x_t, \theta), w_{2t} - f_{2t}(x_t, \theta); \Omega].
 \end{aligned}$$

Estimation of and inference for unknown parameters θ and Ω are carried out by regular means based on equation (3).

The Translog Demand System

For empirical implementation we consider the translog demand system, derived from the indirect utility function (Christensen, Jorgensen, and Lau). Specifically,

$$(4) \quad \log V(p_t, m_t) = \alpha_0 + \sum_{i=1}^n \alpha_i \log(p_{it} / m_t) + \frac{1}{2} \sum_{i=1}^n \sum_{j=1}^n \beta_{ij} \log(p_{it} / m_t) \log(p_{jt} / m_t),$$

where p_t is a vector of prices, m_t is budget, and α_0, α_i and β_{ij} are unknown parameters.

Applying *Roy's identity* to equation (4) yields the translog demand system:

$$(5) \quad w_{it}(p_t, m_t) = \frac{\alpha_i + \sum_j \beta_{ij} \log(p_{jt} / m_t)}{\sum_j \alpha_j + \sum_j \sum_i \beta_{ij} \log(p_{jt} / m_t)}, \quad i = 1, 2, \dots, n,$$

where $w_{it}(p_t, m_t)$ are expenditure shares. Homogeneity is implicit in equations (4) and (5) by use of normalized prices p_{jt} / m_t for all j . To incorporate demographic variables in the demand equations (5), let

$$(6) \quad \alpha_i = \sum_{k=1}^m \alpha_{ik} z_{kt},$$

where z_{1t} is unity. Parametric adding-up and symmetry restrictions are imposed in estimation of

the demand system:

$$(7) \quad \sum_{i=1}^n \alpha_{i1} = -1, \quad \sum_{i=1}^n \alpha_{ik} = 0 \quad \text{for } k \geq 2$$

$$(8) \quad \beta_{ij} = \beta_{ji} \quad \text{for all } i, j.$$

Data

The data for this study are compiled from the USDA's 1987-88 NFCS (U.S. Department of Agriculture). The survey contains quantities and expenditures on more than one hundred different cuts of beef. This study will focus on three disaggregate forms of beef: steaks, roast, and ground beef. Together, these three products constitute over 95% of the beef consumed by households in the United States. Price (unit value) for each product is derived from the reported expenditure and quantity, and regional averages are used as proxies for the non-consuming household residing in each region. Besides prices and total meat expenditure, the explanatory variables include household age composition (numbers of household members in three age groups: age < 20, age 20–64, and age ≥ 65), education of the household head, and dummy variables indicating urbanization (Urban), regions of residence (Northeast, Midwest, South), home ownership (Homeowner), race (White), ethnicity (Hispanic), gender of meal planner (Female planner), and food stamp recipient status (Food stamp).

The NFCS originally contains 4,273 households, but upon recommendation by the USDA about 200 non-housekeeping households are excluded from the sample. In addition, households who did not consume any beef products are excluded. As a result, a total of 3,505 households were selected for the empirical analysis. The sample statistics of all variables are presented in Table 1. The numbers (percentages) of zero observations are 1,974 (51%) for steaks, 2,596 (26%) for roast, and 635 (18%) for ground beef. Therefore, estimation procedures not accounting for such censoring are likely to produce very unreliable results. On average, a

household consumes 1.23 lbs. of steaks, 0.83 lb. of roast, and 1.98 lbs. of ground beef. Among the consuming households, the corresponding averages are 2.53 lbs., 3.20 lbs., and 2.42 lbs. for steaks, roast, and ground beef, respectively.

With respect to sample characteristics, Table 1 shows that a large majority of respondents, 85%, 69%, and 70%, were white, homeowner, and residing in urban area, respectively. Survey respondents with Hispanic origin accounted for only 4% of the sample, while 7% of the respondents were participants of the food stamp program. In addition, 35% of the households resided in the south and 80% of the meal planners were females. Average household size was slightly less than three persons, and the average household head had some college education after finishing high school.

Results

Parameter Estimates

ML estimation is carried out with the ground beef equation deleted, based on the likelihood function (3). Demographic parameters (α_{3k} for all k ; see (6)) and standard deviation (σ_3) for the ground beef equation, along with their standard errors, are calculated using the adding-up restriction (7). The correlation coefficient (ρ_{12}) between steaks and roast is significant at the 1% significance level, which, apart from the need to impose cross-equation restrictions, justifies estimation of the demand functions as a system.

The parameter estimates, presented in Table 2, suggest that all but two of the price parameters (β_{ij}) are statistically significant at the 1% significance level. In general, the three beef products are substitutes for each other and that own price effects are large relative to cross-price effects. These results are to be expected and consistent with findings reported in the literature (Capps and Havlicek; Gao and Spreen; Heien and Pompelli).

Unlike in larger demand systems in which the number of parameters would increase exponentially, the small system considered in the current study allows the inclusion of a relatively large number of demographic variables. Except white, all demographic variables are statistically significant in at least one of the equations, justifying the inclusion of these variables as demand shifters. Over all, the effects of socioeconomic and demographic variables are much smaller in magnitudes than the price effects, except for roast beef in which the own-price effect was not statistically significant. While the results confirm the importance of demographic variables in accommodating heterogeneous preference (and influencing demand), they also indicate that prices are the most important factors, if considered individually, to affect quantity of each beef product demanded.

The effects of demographic characteristics should be interpreted in a relative sense because the estimated coefficients sum to zero across the beef products (see adding-up restrictions (7)). The signs and magnitudes of the estimated coefficients are important as they indicate relative changes in beef consumption across the three beef product categories given a change in the demographic characteristics. Household composition appears to be one of the most important demographic factors that influences demand across all three beef products. The decrease in ground beef consumption due to the addition of a household member younger than 20 years of age reflects a relative shift in the allocation of beef consumption from ground beef to steaks and roast. Whereas an increase in household size (regardless of age group) increases demand for steaks, the demand for steaks decreases as a household member grows older. An addition of a household member younger than 20 years old will increase the share of steaks by 0.086, while adding a person of 65 years or older will only increase steaks consumption by 0.064. In contrast, an increase of household member 65 years or older will decrease household

demand for beef roast, and an increase of household member younger than 65 years of age will also decrease demand for ground beef.

The urbanization and regional location of respondent households also exert significant impacts on beef consumption. Urban households would increase consumption of roast and ground beef at the expense of steaks, while households residing in the Midwest and Southern regions would increase consumption of steaks primarily at the expense of ground beef. These results are consistent with the conventional view that people in the Midwest have a habit of eating more steaks than their counterparts in the rest of the United States. Homeowners tend to consume more ground beef relative to roast and steaks.

Hispanic households have significant influences in the demand for all beef products. The result suggests that Hispanics seem to have stronger preferences toward roast than steaks and ground beef. This result stands in contrast to findings by Heien and Pompelli, who reported that Hispanic households had a significantly positive and negative impact on consumption of steaks and roast, respectively. The income level of Hispanic households is generally lower than that of non-Hispanic households. Viewed in this perspective, the finding that Hispanics would increase and decrease consumption of roast and steaks, respectively, as compared to households with other ethnic backgrounds, seems reasonable as might be expected.

Furthermore, the results show that food stamp participants have a positive influence on the demand for steaks and a negative effect on ground beef. To the extent that food stamps are transfer payment that increases participants' income or food spending power, this result is consistent with what might be expected as an income effect. Interestingly, this finding tends to confirm a popular perception that food stamps afford the recipients to upgrade their food

purchasing with more luxury food items instead of upgrading their diets to be more nutritionally adequate.

Demand Elasticities

Unlike in more “traditional” demand system modeling, censoring in the dependent variables must be accommodated in calculation of the elasticities. Parallel to McDonald and Moffitt for the univariate Tobit model, we investigate the effects of prices, total meat expenditure, and demographic variables by decomposing the dependent variables. For product i , the probability, conditional mean, and unconditional mean are, respectively,

$$(8) \quad \Pr(w_i > 0) = \Phi[f_i(x, \theta)/\sigma_i],$$

$$(9) \quad E(w_i | w_i > 0) = f_i(x, \theta) + \sigma_i \phi[f_i(x, \theta)/\sigma_i] / \Phi[f_i(x, \theta)/\sigma_i],$$

$$(10) \quad E(w_i) = \Pr(w_i > 0) E(w_i | w_i > 0),$$

where $\phi(\cdot)$ and $\Phi(\cdot)$ are univariate standard normal probability density and probability distribution functions, respectively.

The elasticities with respect to prices and total expenditure, along with standard errors for these elasticities (derived by the *delta method*), are presented in Table 3. The total (unconditional) price and expenditure elasticities are decomposed into two components in terms of probability and conditional elasticities. The probability elasticity represents the percentage change in probability that a household will enter or exit the beef market given a percentage change in the price of beef or household expenditure. The conditional elasticity represents the percentage change in the beef consumption level of the consuming households given a percentage change in beef price or household expenditure. The unconditional elasticity, as the sum of probability and conditional elasticities, is an overall measure of responsiveness of quantity demanded to changes in an explanatory variable.

As expected, all own-price elasticities of probability are negative and statistically significantly different from zero at the 1% significance level. Among the six cross-price elasticities or probability, two are not statistically significant and three are negative (including one that is not significant), suggesting a complementary instead of a substitution relationship between steaks and roast. Moreover, it is noted that all the conditional elasticities have the same signs as the probability elasticities. As a result, they reinforce each other in contributing to the total (unconditional) elasticities. All conditional own-price elasticities are negative and greater than unity (in absolute values).

Based on the unconditional elasticities, our results suggest that the demands for the three beef products examined in this study are all price elastic. Our own-price elasticities are much lower than those reported by Capps and Havlicek, which ranged from -1.52 for ground beef to -1.83 for roast. Other analysts, however, found much lower own-price elasticities. For instance, Heien and Pompelli found the own-price elasticities range from -0.73 for steaks to -1.11 for roast, while estimates reported by Gao and Spreen vary from -0.43 for ground beef to -0.65 for roast. In contrast, closer agreements are found among the expenditure elasticities. Consistent with previous studies (Capps and Havlicek; Gao and Spreen; Heien and Pompelli), this study shows that both steaks and roast are expenditure elastic with roast being relatively more responsive to expenditure changes than steaks. Two of the previous studies (Gao and Spreen; Heien and Pompelli) also found that the demand for ground beef was inelastic with respect to household expenditure, while Capps and Havlicek suggested otherwise. On the other hand, Nayga estimated the demand for beef cuts are very income inelastic, ranging from 0.08 for roast to 0.14 for steaks.

While it is extremely difficult to compare elasticity estimates across studies and it is unclear to what extent these different elasticities are caused by the different estimation procedures, the

demand elasticities obtained in this study appear to fall in between the bounds of those reported in the meat demand literature. However, it is important to note that most of these previous findings should be viewed with cautions because, as mentioned above, the methodologies differ greatly among these studies, ranging from use of an inconsistent estimation procedure to complete ignorance of the censoring issues in the sample.

The household composition elasticities are presented in Table 4. The results show that seven of the nine elasticity estimates are statistically significantly different from zero at the 5% significance level or less. In general, the magnitudes of household composition elasticities are fairly small, suggesting that the demand for beef products is relatively unresponsive to changes in household members in a certain age group.

The results show that household composition elasticities range from -0.33 for steaks to 0.42 for roast. Although not strictly comparable, Nayga also indicated that the demand for beef is irresponsive to changes in household size. His estimates of household size elasticities ranged from 0.37 for steak to 0.70 for roast. It is interesting to note that changes in household composition generally have greater impacts on the demand for beef products in terms of changes in probabilities than changes in consumption level. In other words, changes in beef consumption due to changes in household composition are more likely to be attributed to entry by households into or exit from the marketplace than increases or decreases of the amount of beef consumed. This result appears reasonable in the sense that an increase or decrease in household size is more likely to influence the probability of purchasing beef for household consumption than consuming the beef per se.

Summary and Conclusions

This study demonstrated the estimation of a system of demand equations for disaggregated beef products based on the USDA 1987-88 NFCS. A full-information ML estimator applicable to a

censored system of equations was developed. The system of demand equations for steaks, roast and ground beef was specified based on the translog functional form, and demand elasticities with respect to price, expenditure, and household composition were estimated. The current study is different from most previous studies in at least two important aspects: (i) application of the more efficient ML estimation procedure (than two-step estimates); (ii) decomposition of the demand elasticities into probability and conditional elasticities to accommodate censoring in the dependent variables.

In general, the estimated coefficients are highly significant for most of the price as well as demographic variables. The results suggest that the demands for the three beef cuts examined in this study are price elastic. The demands for steaks and roast are also elastic with respect to household expenditure. Although prices appear to be the dominating factors influencing the demand for beef products, the results also document the important and significant effects of demographic characteristics, such as household composition, urbanization, regional location, homeownership, ethnicity, gender of meal planner, and food stamp participation, on demand for beef products.

The decomposition of the total elasticity into probability and conditional elasticities also serves as an important means of providing additional insights to assess the effects of a given change in the price or demographic variable on the probability of consumption and the level of consumption, respectively. As might be expected, the results show that the level of consumption in different beef cuts are mostly affected by a change in beef price or household expenditure. In contrast, the effects of changes in demographic profile, such as household composition, on the demand for beef are reflected mostly in the changes of the consumption probability through household's entry or exit of the beef market.

Although the NFCS data set used in the current study is old and dated, it is nevertheless the most recent among existing data sets that contain price information. The findings confirm the importance of demographic influences on the demand for beef, and show that more insights about the nature of demand for beef can be gained by examining the demand for individual cuts instead of beef in general. Some important marketing implications could be drawn from the results of this study. With the emergence of the Hispanic population in the United States as an important ethnic market and a major sector of the U.S. economy, the finding of Hispanic households' beef-eating habits and preferences provides useful insight and information for the beef industry to develop its marketing strategies that focus on this growing market segment. The finding that food stamps recipients tend to purchase more steaks at the expense of ground beef could be an interesting issue on the assessment of the effectiveness the food assistance program in improving the nutritional status of the recipients.

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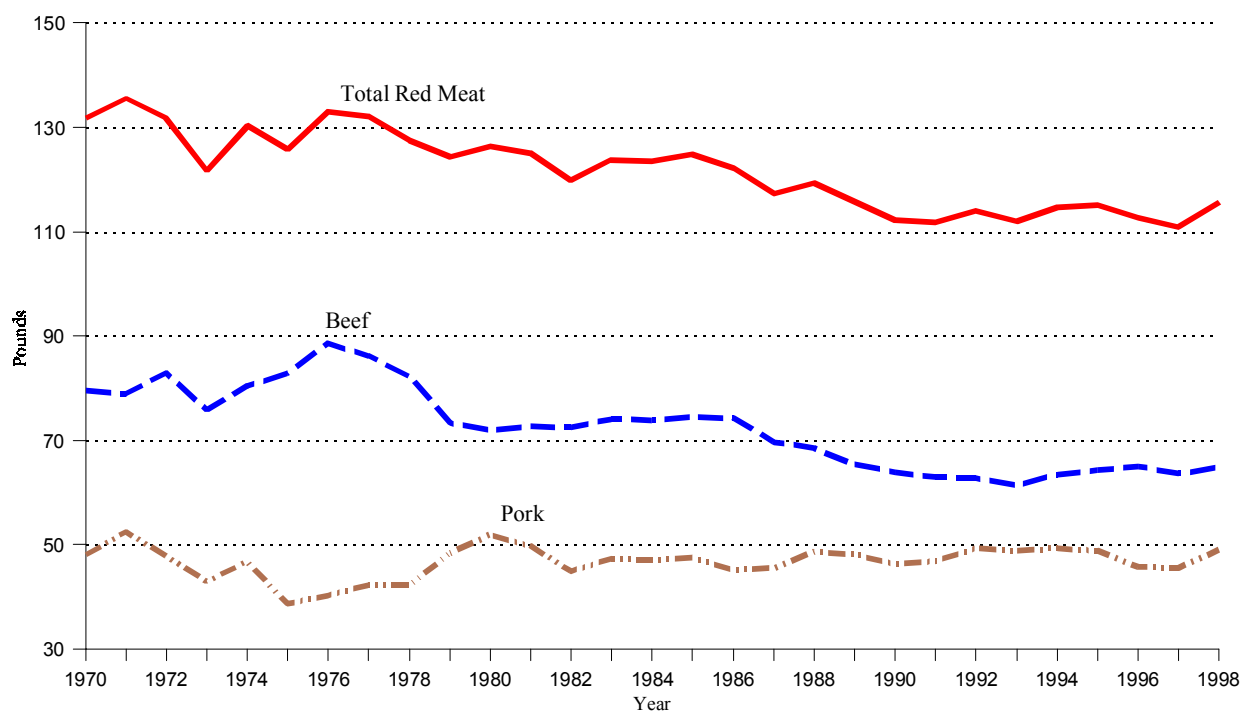
Figure 1. U.S. Per Capita Red Meat Consumption, 1970-1998.

Table 1. Sample Statistics

Variables	Mean	Standard Deviation
Quantities (lb.)		
Steaks: Full sample	1.23	2.02
Consuming households (1,711 observations)	2.53	2.25
Roast: Full sample	0.83	1.66
Consuming households (909 observations)	3.20	1.74
Ground: Full sample	1.98	1.93
Consuming households (2,870 observations)	2.42	1.87
Prices (\$ / lb.)		
Steaks	2.67	0.98
Roast	1.85	0.38
Ground	1.52	0.43
Expenditure shares		
Steaks	0.32	0.37
Roast	0.15	0.29
Ground	0.53	0.40
Demographic variables:		
Age < 20 (number of members)	0.99	1.23
Age 20–64	1.64	0.92
Age ≥ 65	0.31	0.62
Education (years of; household head):	2.36	0.80
= 1, if years of education ≤ 8;		
= 2, if 9 ≤ years of education ≤ 12;		
= 3, if 13 ≤ years of education ≤ 16;		
= 4, if years of education ≥ 17		
Demographic variables (binary)		
Urban (Household resides in)	0.70	0.46
Northeast	0.20	0.40
Midwest	0.28	0.45
South	0.35	0.48
Homeowner	0.69	0.46
White	0.85	0.35
Hispanic	0.04	0.20
Female (meal) planner	0.80	0.40
Food stamp (recipient)	0.07	0.26

Source: Compiled from the Nationwide Food Consumption Survey, 1987-88. Sample size = 3,505.

Table 2. ML Estimation of Translog Demand System with Censored Dependent Variables

Variable	Beef products		
	Steaks	Roast	Ground
Constant	-0.021 (0.063)	0.773*** (0.093)	-1.753*** (0.085)
Price of steaks	0.209*** (0.025)		
Price of roast	0.130*** (0.032)	-0.032 (0.048)	
Price of ground	-0.023 (0.025)	0.250*** (0.034)	-0.757*** (0.032)
Demographic variables			
Age < 20	0.086*** (0.010)	0.047*** (0.013)	-0.134*** (0.012)
Age 20–64	0.072*** (0.015)	0.027 (0.021)	-0.099*** (0.019)
Age ≥ 65	0.064*** (0.022)	-0.068** (0.032)	0.004 (0.027)
Education	-0.019 (0.013)	-0.010 (0.019)	0.029* (0.017)
Urban	-0.083*** (0.022)	0.067** (0.031)	0.016 (0.027)
Northeast	0.018 (0.032)	0.006 (0.048)	-0.024 (0.042)
Midwest	0.141*** (0.030)	-0.016 (0.044)	-0.125*** (0.039)
South	0.072** (0.029)	-0.031 (0.042)	-0.041 (0.037)
Homeowner	-0.011 (0.023)	-0.069** (0.034)	0.079*** (0.030)
White	0.009 (0.030)	0.013 (0.042)	-0.023 (0.037)
Hispanic	-0.122** (0.052)	0.290*** (0.085)	-0.167** (0.076)
Female planner	-0.009 (0.025)	-0.074** (0.037)	0.082** (0.033)
Food stamp	0.126*** (0.042)	0.009 (0.059)	-0.135*** (0.053)
Standard deviation (σ)	0.600*** (0.011)	0.750*** (0.020)	0.483*** (0.015)
Correlation ($\rho_{\text{steaks, roast}}$)	-0.766*** (0.013)		
Log-likelihood	-4,274.994		

Note: Asymptotic standard errors in parentheses. Statistical significance: *** = 1%, ** = 5%, * = 10%. All coefficients of demographic variables ($\alpha_{3k} \forall k$) and standard deviation (σ_3) for ground beef are derived using the adding-up restrictions.

Table 3. Price and Expenditure Elasticities

Quantity	Prices of			Expenditure
	Steaks	Roast	Ground	
	Probability			
Steaks	-0.175*** (0.055)	-0.082 (0.051)	-0.067 (0.043)	0.325*** (0.052)
Roast	-0.491*** (0.120)	-0.238** (0.116)	0.014 (0.109)	0.715*** (0.100)
Ground	0.002 (0.001)	0.001* (0.001)	-0.000 (0.000)	-0.003 (0.002)
	Conditional level			
Steaks	-1.129*** (0.028)	-0.061* (0.032)	-0.050 (0.037)	1.240*** (0.011)
Roast	-0.160*** (0.024)	-1.077*** (0.032)	0.005 (0.035)	1.232*** (0.010)
Ground	0.340*** (0.018)	0.194*** (0.023)	-1.050*** (0.036)	0.516*** (0.021)
	Unconditional level			
Steaks	-1.304*** (0.082)	-0.143* (0.083)	-0.117 (0.080)	1.565*** (0.060)
Roast	-0.651*** (0.142)	-1.315*** (0.147)	0.019 (0.144)	1.947*** (0.103)
Ground	0.342*** (0.018)	0.195*** (0.023)	-1.050*** (0.036)	0.513*** (0.022)

Note: Asymptotic standard errors in parentheses. Statistical significance: *** = 1%, ** = 5%, and * = 10%.

Table 4. Elasticities with respect to Household Age Composition

Quantity	Age < 20	Age 20–64	Age ≥ 65
	Probability		
Steaks	–0.079*** (0.014)	–0.157*** (0.044)	–0.189** (0.082)
Roast	–0.068*** (0.021)	–0.094 (0.078)	0.316*** (0.128)
Ground	0.001* (0.000)	0.001 (0.001)	0.000 (0.000)
	Conditional level		
Steaks	–0.059*** (0.006)	–0.116*** (0.023)	–0.140*** (0.047)
Roast	–0.022*** (0.006)	–0.030 (0.023)	0.103** (0.050)
Ground	0.073*** (0.007)	0.129*** (0.022)	–0.007 (0.048)
	Unconditional level		
Steaks	–0.138*** (0.018)	–0.273*** (0.066)	–0.329*** (0.128)
Roast	–0.090*** (0.027)	–0.124 (0.101)	0.419** (0.177)
Ground	0.073*** (0.007)	0.130*** (0.022)	–0.007 (0.048)

NOTE: Asymptotic standard errors in parentheses. Statistical significance: *** = 1%, ** = 5%, * = 10%.