Estimating Cross Elasticities of Demand for Beef

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This paper examines the extent to which observed changes in per capita beef consumption can be accounted for by changes in relative prices and per capita real income. Particular attention was given to specifying the functional form by focusing on the nature of demand shifts over time. These considerations led to selection of the Fourier flexible form with a Fourier series expansion in relative meat prices. The results support the hypothesis that recent shifts in demand for beef can be attributed to changes in relative prices of competing meats, especially poultry.

Considerable interest has been expressed in the cause of the decline in U.S. beef consumption in recent years. Two recent studies [Chavas; Moschini and Meilke] indicate different reasons for the shift in demand. Chavas attributes changed elasticities since 1970 to increased nutritional consciousness of consumers, while Moschini and Meilke attribute the recent decline in beef consumption to changes in relative prices and real income. Both studies, however, potentially suffer from specification bias. Chavas utilizes a constant elasticity functional form in his base model, although there is no compelling reason to expect elasticities to remain constant over time. Moschini and Meilke attempt to overcome functional form bias by employing a Box-Cox transformation; however, they do not include a separate price variable for poultry in their demand specifications. In both studies, fish prices are excluded from the demand specifications.

In this study, a more general approach is taken to specifying and estimating demand for beef. The maintained hypothesis is that changes in beef consumption over time can be explained by changes in relative prices and real income. Demand for beef is specified to depend upon poultry and fish prices as well as beef and pork prices and income. In order to avoid functional form misspecification, the seminonparametric methodology of the Fourier flexible form is utilized. The results support the hypothesis that recent shifts in demand for beef can be attributed to changes in relative prices of competing meats, especially poultry.

Nature of the Data

Data used in demand analysis of beef are annual time series on per capita beef and veal consumption and retail price indexes for beef and veal, pork, poultry, and fish. Consumer income is measured by per capita total personal consumption expenditures. All price and income data are deflated by the consumer price index. The analysis covers the period 1947–83. Data sources are USDA, *Food Consumption, Prices, and Expenditures* for quantity and price data and the *Economic Report of*

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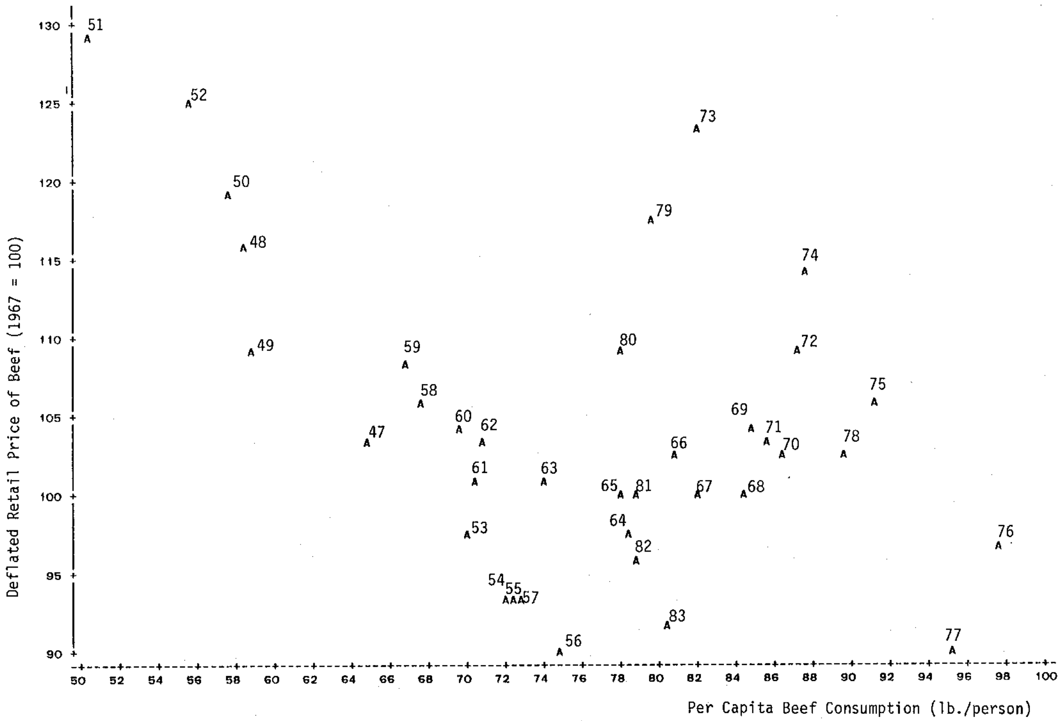

Figure 1. Per Capita Consumption and Deflated Retail Price of Beef, 1947–83.

the President for data on personal consumption expenditures and the consumer price index.

Figure 1 shows the relationship over time between deflated beef price and per capita beef consumption. Each pair of observations corresponds to a particular year. This diagram clearly suggests that the demand curve for beef has shifted over time. These data are grouped into four sets of observations: one set of observations for the years 1947 through 1957, another for the years 1958 through 1968, a third set for the years 1969 through 1979, and the fourth set for the years 1980 through 1983. These clusters of points give a rough indication of the nature of the shifts in the demand curve for beef over time, although within each group there are clearly additional shifts in the demand curve. These data groupings, particularly the first grouping for 1947 through 1957, indicate that the relationship between price and quantity along a given demand curve has a curvilinear rather than a linear relationship. Thus, a logarithmic transformation would appear to be appropriate.

How do we account for the shifts in the demand curve for beef over time? Increases in demand for beef until the early 1970s appear to be attributable mainly to the growth in real income. This can be seen in plots of the residuals for log price and log quantity obtained by regressing log price on log income and log quantity on log income, respectively. The importance of income as a demand shifter also can be ascertained from the regression of log per capita beef quantity (LQB) on log deflated beef price (LDPB), log deflated pork price (LDPPK), log deflated poultry price (LDPPY), log deflated fish price (LDPF), and log deflated per capita total personal consumption expenditures (LDY). For the years 1947 through 1972, the estimated regression equation is

$$\begin{aligned} \hat{LQB} = {} & \underset{(0.649)}{6.678} - \underset{(0.075)}{0.834}LDPB \\ & + \underset{(0.074)}{0.176}LDPPK + \underset{(0.058)}{0.023}LDPPY \\ & + \underset{(0.150)}{0.128}LDPF + \underset{(0.094)}{0.730}LDY, \\ & \bar{R}^2 = 0.979, \quad DW = 1.73, \end{aligned} \tag{1}$$

where the values in parentheses are estimated standard errors of the regression coefficients.

In contrast to the earlier period, shifts in the demand curve for beef since the early 1970s appear to be attributable mainly to changes in relative prices of competing meats. From 1973 through 1983, real income as measured by deflated per capita total personal consumption expenditures only increased about seven percent. While real beef prices declined about 25 percent over this period, real pork prices declined about 30 percent, and real poultry prices declined almost 45 percent. Real fish prices cycled over this period, increasing about 15 percent between 1973 and 1978 and then declining by about the same amount between 1978 and 1983. The fact that real prices of competing meat products—particularly poultry—have declined so much in the last decade provides the motivation for focusing attention on cross elasticities of demand for beef.

Model Specification

Statistical analysis of demand for beef is conducted within the context of a single equation framework. Per capita demand for beef depends on a number of factors including (a) per capita income; (b) retail prices of beef, pork, poultry, and fish; and (c) retail prices of nonmeat consumer goods. Effects of prices of nonmeat goods are assumed to be captured in an aggregate price index which moves proportionately over time with the consumer price index (CPI). By the zero homogeneity property of consumer demand functions, the CPI can be used as a deflator for income and retail meat prices so that per capita beef demand is related to per capita real income and real retail prices of beef, pork, poultry, and fish. The problem statistically is then selection of a functional form which best fits the data.

Consider as one candidate for a functional form the constant elasticity form in which the logarithm of per capita demand is linearly related to logarithms of real prices and the logarithm of per capita real income. Equation (1) indicates that this functional form provides a satisfactory fit to the data over the period 1947 through 1972. Since the early 1970s, however, the nature of the demand shifts have been such as to make the assumption of constant elasticities over the entire sample period suspect. This is evidenced by changed elasticities—especially cross elasticities—and a low Durbin-Watson statistic when this functional form is fitted to the data over the period 1947 through 1983:[1]

$$\begin{aligned} \hat{LQB} = {} & \underset{(0.457)}{5.950} - \underset{(0.087)}{0.600}LDPB \\ & + \underset{(0.103)}{0.357}LDPPK + \underset{(0.094)}{0.117}LDPPY \\ & - \underset{(0.086)}{0.396}LDPF + \underset{(0.144)}{0.955}LDY, \\ & \bar{R}^2 = 0.927, \quad DW = 1.05. \end{aligned} \tag{2}$$

These considerations indicate a logical extension of equation (2) is to make the cross elasticity coefficients depend on real prices of the competing meat products. One possible specification of this behavior is

$$\begin{aligned} LQB = {} & A_0 + A_1LDPB + A_2LDPPK + A_3LDPPY \\ & + A_4LDPF + A_5LDY \\ & + A_6LDPB \times LDPPK \\ & + A_7LDPB \times LDPPY \end{aligned}$$

[1] Equality of elasticities between the two subperiods 1947–72 and 1973–83 was tested using the Chow test based on the F-distribution. The null hypothesis of equality of elasticities between the subperiods was rejected at a probability level smaller than one percent. See Fisher for a concise presentation of the Chow tests for structural change.

$$+ A_8 LDPB \times LDPF + A_9 LDPPK \times LDPPY + A_{10} LDPPK \times LDPF + A_{11} LDPPY \times LDPF + U_1, \quad (3)$$

where U_1 is a random disturbance term. This specification is similar to that used by Waugh [p. 17] in estimating the cross elasticity of demand for food. The model could be expanded to include squared terms involving log prices and log income as well as interaction terms with log income. But the statistical results indicated these terms collectively would provide little additional explanation.

A more general functional form—one that is not conditional on any particular parametric functional form—is the Fourier flexible form described by Gallant [1984]. This functional form has the property of Sobolev flexibility. This means the error of approximation of a functional form and its derivatives up to some specified order can be made arbitrarily small by including additional terms in the Fourier series as the sample size increases. The Fourier form also can estimate elasticities consistently [El Badawi *et al.*].

The multivariate Fourier series can be written as

$$M(x) = \text{constant} + 2 \sum_{\alpha=1}^{A} \sum_{j=1}^{J} \cdot [U_{j\alpha}\cos(jk'_\alpha x) - V_{j\alpha}\sin(jk'_\alpha x)],$$

where x is the vector of logarithms of real prices and real per capita income. The k'_αs are multi-indexes which indicate the direction of the Fourier series expansion. Rules for obtaining these indexes are given in Gallant [1981, sec. 2]. By adding linear and quadratic terms, one obtains the Fourier flexible form [Gallant, 1984].

The particular specification of the Fourier form used in this study consists of the linear and quadratic terms in (3) plus the multivariate Fourier series with the six multi-indexes: $k_1 = (1, -1, 0, 0, 0)'$, $k_2 = (1, 0, -1, 0, 0)'$, $k_3 = (1, 0, 0, -1, 0)'$, $k_4 = (0, 1, -1, 0, 0)'$, $k_5 = (0, 1, 0, -1, 0)'$, and $k_6 = (0, 0, 1, -1, 0)'$. Here x_1, x_2, x_3, and x_4 denote the logarithms of real prices for beef, pork, poultry, and fish respectively; x_5 denotes the logarithm of real per capita income. Letting $f(x_1, x_2, x_3, x_4, x_5) = f(x)$ denote the deterministic component of (3), the Fourier specification used here, therefore, can be thought of as equation (3) plus a multivariate Fourier series expansion in relative meat prices; that is,

$$\begin{aligned} LQB = f(x) + 2 \sum_{j=1}^{J} [& U_{j1}\cos(j(x_1 - x_2)) - V_{j1}\sin(j(x_1 - x_2)) \\ & + U_{j2}\cos(j(x_1 - x_3)) - V_{j2}\sin(j(x_1 - x_3)) \\ & + U_{j3}\cos(j(x_1 - x_4)) - V_{j3}\sin(j(x_1 - x_4)) \\ & + U_{j4}\cos(j(x_2 - x_3)) - V_{j4}\sin(j(x_2 - x_3)) \\ & + U_{j5}\cos(j(x_2 - x_4)) - V_{j5}\sin(j(x_2 - x_4)) \\ & + U_{j6}\cos(j(x_3 - x_4)) - V_{j6}\sin(j(x_3 - x_4))] + U_2, \end{aligned} \quad (4)$$

where U_2 is a random disturbance term. Note that this form of the Fourier form differs from that indicated by Gallant [1984, p. 207] in that the quadratic terms in the two specifications differ. This difference has no effect on the approximation properties of the Fourier form since approximation properties are based on the Fourier series [Gallant, 1981].

With a Fourier series approximation, one must scale the data so that they lie in the interval $(0, 2\pi)$. This is because a Fourier series approximation near a point of discontinuity can oscillate wildly [Gallant, 1984, p. 206]. To avoid this problem, the meat price data were rescaled as follows:

$$x_1 = LDPB = LOG(DPB) + 1.1418,$$
$$x_2 = LDPPK = LOG(DPPK) + 1.20019,$$
$$x_3 = LDPPY = LOG(DPPY) + 0.64432,$$
$$x_4 = LDPF = LOG(DPF) + 1.05165,$$

where DPB, DPPK, DPPY, and DPF denote deflated retail price indexes for beef and veal, pork, poultry, and fish, respectively. This scaling ensures that each price series is between 0 and 6, and that the difference in any two price series is never zero.

Statistical Results

Equations (3) and (4) were estimated by ordinary least squares with data over the period 1947 through 1983. Initial estimations indicated that the interaction terms with the price of pork added little to the explanation of beef consumption. Thus, these terms were deleted from each model.[2]

The results for equations (3) and (4) are presented in Table 1 as models A and C. The results for model C are for J = 1. Both models have high explanatory power as indicated by the adjusted R-squared values. In both models, highly significant interaction terms between beef and poultry, beef and fish, and poultry and fish are indicated. The Fourier form (model C), however, is superior to the quadratic specification. The F-test that the sine/cosine terms are zero is 8.40. With 6 numerator and 22 demoninator degrees of freedom, one would reject the quadratic specification at a significance level smaller than one percent.

In an attempt to determine whether the difference in explanatory power of the two models can be attributed to structural change, the quadratic specification was subjected to two different Chow tests: one for the subperiods 1947–72/1973–83 and the other for the subperiods 1947–79/1980–83. These breaks in the sample were suggested by the scatter diagram shown in Figure 1. The tests indicated significant parameter change between the first two subperiods (1947–72/1973–83) but insignificant change between the second two subperiods (1947–79/1980–83).

Econometric results for the quadratic demand model allowing for parameter change between 1947–72 and 1973–83 are presented as model B in Table 1. (The variable D is the dummy variable which equals one for 1973–83 and zero otherwise.) Based on the adjusted R-squared values for models B and C, which are quite close, one might be tempted to conclude that structural change is indistinguishable from functional form misspecification. This inference, however, is incorrect as the following statistical analysis reveals.

The relative explanatory power of models B and C can be compared using the J-test for nonnested regression models proposed by Davidson and MacKinnon. The J-test is conducted by estimating a compound regression model which consists of the regressors from one model and predicted values of the dependent variable from the other model, and then testing the significance of the predicted variable using a conventional t-test. Under the null hypothesis that model B is correct, the t-value for the predicted values from model C is 3.6. Under the null hypothesis that model C is correct, the t-value for the predicted values from model B is −0.4. The interpretation of these tests is that the truth of B is rejected, while the truth of model C (the Fourier form) cannot be rejected. Since the Fourier model assumes parameter constancy, this suggests significant parameter change in the quadratic specification between 1947–72 and 1973–83 can be attributed to functional form misspecification. In other words, the results are consistent with the hypothesis that shifts in demand for beef can be attributed to changes in real meat prices and real income.

Elasticities

Price and income elasticities for the three models for selected years are presented in Table 2. Not surprisingly, the elasticities vary somewhat from one model to the other. Yet, the pattern of change in elasticities over time is remarkably similar among models. In particular, all three

[2] F-values for equations (3) and (4) with interaction terms for the price of pork were both insignificant at the five percent probability level.

TABLE 1. Econometric Estimates of U.S. Beef Demand, 1947–83.

Explanatory Variable, Statistic	Model (Dependent Variable LQB)		
	A	B	C
Constant	−74.600 (31.947)	−86.733 (34.443)	−78.117 (31.210)
LDPB	18.872 (5.293)	10.479 (5.864)	23.897 (8.495)
LDPPK	0.173 (0.074)	0.149 (0.058)	0.164 (0.052)
LDPPY	1.993 (2.083)	5.133 (2.070)	6.626 (2.994)
LDPF	6.400 (4.399)	16.707 (6.968)	−1.263 (9.654)
LDY	0.796 (0.102)	0.719 (0.083)	0.681 (0.081)
LDPB × LDPPY	−1.373 (0.254)	0.027 (0.403)	−0.089 (0.572)
LDPB × LDPF	−2.094 (0.708)	−2.008 (1.097)	−1.500 (0.714)
LDPPY × LDPF	1.066 (0.269)	−0.929 (0.439)	−1.048 (0.867)
D × LDPPY		15.483 (3.791)	
D × LDPF		−13.597 (3.416)	
D × LDPB × LDPPY		−2.503 (0.618)	
D × LDPB × LDPF		2.249 (0.584)	
D × LDPPY × LDPF		−0.058 (0.047)	
sin(LDPB-LDPPY)			−4.643 (1.634)
cos(LDPB-LDPPY)			0.627 (0.896)
sin(LDPB-LDPF)			−11.271 (6.262)
cos(LDPB-LDPF)			−0.489 (0.982)
sin(LDPPY-LDPF)			−4.873 (0.770)
cos(LDPPY-LDPF)			−0.211 (0.614)
$\bar{R}^2$	0.971	0.986	0.989
SSE	0.01909	0.00780	0.00580
D.W.	2.24	2.45	1.88

models indicate that since about the middle 1970s the cross price elasticity with respect to poultry has increased substantially, and that the relationship between beef and fish has changed from one of substitutability to one of complementarity.

The finding that beef demand has be-

TABLE 2. Estimated Elasticities of U.S. Beef Demand for Alternative Models, Selected Years.

Model	Year	Elasticity with Respect to Price of				Income Elasticity
		Beef	Pork	Poultry	Fish	
A	1950	−1.14	0.17	−0.09	0.32	0.80
	1960	−0.35	0.17	0.03	0.08	0.80
	1965	−0.19	0.17	0.09	0.04	0.80
	1970	−0.11	0.17	0.11	−0.17	0.80
	1975	−0.68	0.17	0.31	−0.14	0.80
	1978	−0.73	0.17	0.46	−0.23	0.80
	1980	−0.44	0.17	0.32	−0.50	0.80
B	1950	−0.74	0.15	0.03	−0.70	0.72
	1960	−0.64	0.15	0.07	0.04	0.72
	1965	−0.66	0.15	0.07	0.23	0.72
	1970	−0.77	0.15	0.02	0.30	0.72
	1975	−1.12	0.15	0.44	−0.68	0.72
	1978	−0.76	0.15	0.39	−0.56	0.72
	1980	−0.45	0.15	0.29	−0.40	0.72
C	1950	−0.57	0.16	0.13	−0.90	0.68
	1960	−0.61	0.16	−0.03	0.22	0.68
	1965	−0.60	0.16	−0.01	0.37	0.68
	1970	−0.43	0.16	−0.03	0.20	0.68
	1975	−0.97	0.16	0.18	−0.15	0.68
	1978	−0.74	0.16	0.61	−0.91	0.68
	1980	−0.27	0.16	0.44	−1.02	0.68

come more sensitive to poultry prices in recent years is intuitive, yet new. Previous studies apparently have failed to uncover this relationship either because of functional form misspecification or because of lack of recent data which show this relationship. An apparent U-shaped relationship of this cross elasticity with respect to time might suggest why the constant elasticity model, equation (2), fails to show a significant relationship between beef demand and poultry price.

The changed relationship between beef and fish from substitutability to complementarity seems counter-intuitive and perhaps indicates spurious correlation. However, this changed elasticity relationship might reflect a desire on the part of consumers, in response to health concerns about red meat, to expand their diets to include fish.

Own-price elasticities vary from one model to another and over time within each model. However, there is little evidence to indicate own elasticities have been increasing or decreasing in recent years.

Cross-price elasticities with respect to pork and income elasticities were constrained to be constant in each model. As indicated previously, there was little evidence to indicate these elasticities had changed over time.

To highlight the importance of changes in relative meat prices in accounting for declines in beef demand in recent years, average elasticities from model C for the period 1975 through 1980 are applied to percentage changes in real meat prices and per capita real income over this period. For this period per capita beef consumption declined 15 percent, real beef prices increased four percent, real pork prices declined 31 percent, real poultry prices declined 24 percent, real fish prices increased six percent, and per capita real income increased six percent. Average elasticities of beef demand over this period were −0.64, 0.16, 0.49, −0.79, and 0.68 for beef, pork, poultry, fish, and income, respectively. Multiplying elasticities times percentage changes and summing yields a predicted decline in per capita consumption of 20 percent. Of this predicted decline in consumption, almost 12 percent can be ascribed to decreased poultry prices, five percent to decreased pork prices, and five percent to increased fish prices. Increased beef prices accounted for less than three percent of the decline in beef consumption over this period. Other factors constant, increases in real income would have increased demand for beef by only four percent.

Conclusions

The purpose of this study was to determine the extent to which observed changes in per capita beef consumption can be accounted for by changes in relative prices and per capital real income. Particular attention was given to specifying the functional form by focusing on the nature of the demand shifts over time. These considerations led to selection of the Fourier flexible form with a Fourier series expansion in relative meat prices. The results are consistent with the stated hypothesis and lend support to the contention that changed elasticities over time are consistent with a stable demand structure.

A significant finding of this study is that, since about the middle 1970s, beef demand has become more sensitive to poultry prices. This suggests that the beef industry should pay close attention to developments in the poultry industry. In particular, further reductions in poultry costs, if not matched by decreases in beef production costs, can have strong adverse effects on the demand for beef. It, therefore, would seem advantageous for the industry to seek ways to further reduce beef production costs to remain competitive with poultry.

One unresolved problem is the cause of the changed relationship between beef and fish from substitutability to complementarity. Since fish accounts for a very small proportion of the typical consumer's total expenditures, the change in sign of the cross-price elasticity from positive to negative cannot be ascribed to the income effect outweighing the substitution effect. Whether this result is a statistical aberration or not can be resolved only by adopting a system-wide approach in which all demand relationships are estimated jointly with symmetry imposed on the cross-price relationships.

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