Parameter Stability and the U.S. Demand for Beef

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The main objective of this paper is to test the hypothesis that consumer preferences for beef in the United States have been affected by structural change, which reduces to testing for parameter stability in estimated demand equations. To this end, alternative specifications of the demand function are estimated using a general form of the Box-Cox transformation. Tests based on recursive residuals and on the F distribution provide little evidence of structural change, and suggest that the recent decline in beef consumption may be explained by changes in relative prices.

The sizable decline in U.S. beef consumption that has taken place in recent years has led to some speculation, in both the popular and professional literature, that the demand for beef may have been affected by a structural change resulting, ceteris paribus, in consumption levels lower than those of the early 1970s (Bertin; Hieronymus; Chavas). The cause of this alleged structural shift is often ascribed to the increased nutritional consciousness of consumers concerned with limiting their intake of fat and cholesterol.

Whether the decrease in the consumption of beef is a result of changed market conditions (relative prices and income), or reflects a more fundamental change in the underlying consumers' preferences, is an important question for both the beef industry and agricultural economists. Indeed, if the consumption decline is due to market forces, there is little that the beef industry can do, other than wait for a more favorable economic climate. If, however, there has been a structural shift in the demand for beef, the industry needs to pursue the question of why preference relations have changed; and, presumably, undertake corrective measures—such as advertising, consumer educational programs, grading changes, etc.—which could shift preferences back towards their original position.

From the agricultural economist's point of view, the entertained structural shift poses some interesting issues, especially those connected with the choice of an appropriate model for estimation and forecasting purposes. The hypothesis of structural change, in fact, affects the very heart of econometric modeling. For most estimation procedures, a necessary condition is that of constancy of the economic structure generating the sample observations, implying that there exists a single parameter vector relating dependent and independent variables, a single functional form, and a constant set of error process parameters. Since econometrics typically deals with data that have not been generated by controlled experiments, the constancy condition is usually dealt with by assumption. Under structural change, this assumption is obviously unjustified and, indeed, suggestions have been made to pursue the estimation task within a more general framework, namely that of varying parameter models (Rausser *et al.*).

However, the constant parameter mod-

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This project was funded by the Ontario Ministry of Agriculture and Food.

Western Journal of Agricultural Economics, 9(2): 271–282

el is still widely used in empirical work because of its simplicity of estimation, analysis, and interpretation. Consequently, at the empirical level, the problem of ascertaining whether or not the underlying economic structure can safely be assumed constant is of considerable interest. In our case, this problem reduces to that of testing for parameter shifts in the demand for beef, where the notion of parameter should be taken in the broad sense of defining the functional form, the quantitative relationship between dependent and independent variables, and the form of the stochastic process associated with the demand function.

There are, nonetheless, some conceptual problems in trying to detect preference changes by testing the stability of the parameters of estimated demand equations, because parameter instability in econometric models can arise from all sorts of misspecifications (omitted variables, incorrect functional form, etc.). Consequently, only a model devoid of any other type of specification error, i.e. the true model, can properly ascribe the source of parameter shifts to changed preferences. To minimize this problem, a general form of the Box-Cox transformation and several different specifications of the demand for beef are used.[1]

Model Specification

Economic theory suggests that all prices and income enter demand functions. Empirically, however, it is not possible to estimate functions with this degree of generality, and only the variables judged most relevant are included individually, with all other variables captured in an aggregate price index. In this study, per capita demand for beef (PCDBF) is specified as dependent upon the retail price of beef (RPBF), the retail price of pork (RPPK), a price index for all other relevant goods, and an income variable.

In particular, in model A the consumer price index (CPI) is used as the price indicator for all other goods, and the income term is per capita disposable income (PCDY). Model B incorporates a notion of separability of preferences. This assumption implies that commodities can be partitioned into groups so that preferences within groups can be described independently of other groups, or equivalently, that utility maximization entails a two-stage budgeting process (Deaton and Muellbauer). Thus in model B, besides the price of beef and pork, the explanatory variables include a price index of food other than beef and pork (CPIOF), and per capita expenditure on food (PCEXF). Models C and D incorporate the theoretical restriction that demand is homogeneous of degree zero in prices and income. This is achieved by deflating the price and income terms of models A and B. In particular, in model C beef price, pork price and per capita disposable income are deflated by the CPI, whereas in model D beef price, pork price and per capita expenditure on food are deflated by CPIOF.

The introduction of the concept of separability implies that a change in consumer preferences may affect only the first stage of the budgeting process, which suggests that the allocation of income to the food group should be considered explicitly. This is the reason for model E, where per capita expenditure on food is explained by a price index of food (CPIF), a price index of non-food goods (CPINF), and per capita disposable income. Finally, since the data is quarterly, the specification of all five models includes seasonal dummies.

[1] All of the analysis is conducted within a static, single-equation approach. The alternative of a more general model, either dynamic and/or system-wide, was not pursued since testing for parameter instability in these cases is much more difficult. For instance, recursive analysis is possible only on single-equation linear econometric models with nonstochastic regressors.

Model Estimation

As already discussed, it is important that functional form misspecification be avoided. For this reason, a flexible functional form specification based on the Box-Cox transformation is used. This transformation, proposed by Box and Cox and further developed by Zarembka is of the type

$$y_t^{(\lambda)} = \begin{cases} (y_t^\lambda - 1)/\lambda, & \lambda \neq 0 \\ \ln y_t & , \lambda = 0 \end{cases} \tag{1}$$

where the value of λ, a parameter to be estimated, determines the shape of the function. This power transformation, which has been used to estimate meat demand equations previously (Chang; Hassan and Johnson, 1979a; Pope *et al.*), may however introduce unwanted restrictions on price and income elasticities, particularly when both dependent and independent variables are transformed by the same parameter (Gemmill). Consequently, it is desirable to allow for different transformation parameters for different variables. Three transformation parameters are therefore used: λ for the dependent variable, μ for the price variables, and θ for the income term. Thus, letting y represent the dependent variable, X_1 the matrix of constant and dummy variables, X_2 the matrix of price variables, and X_3 the income variable, models A to E can be represented as

$$y^{(\lambda)} = X_1\beta_1 + X_2^{(\mu)}\beta_2 + X_3^{(\theta)}\beta_3 + u. \tag{2}$$

By assuming that u is a vector of independently distributed normal variates,[2] this model can be estimated using a maximum likelihood approach. Although widely used, this approach can lead to serious bias in estimation and hypothesis testing if the disturbances are in fact serially correlated, as already illustrated for the case of beef by Blaylock and Smallwood. Thus, following Savin and White, it is assumed that the residuals of (2) follow a first-order autoregressive process of the type $u_t = e_t + \rho u_{t-1}$. For this stochastic specification, the concentrated log-likelihood function of equation (2) is, apart from a constant,

$$L(\lambda, \mu, \theta, \rho) = -\frac{T}{2} \ln \hat{\sigma}^2(\lambda, \mu, \theta, \rho) + \frac{1}{2} \ln(1 - \rho^2) + (\lambda - 1) \sum_{t=1}^{T} \ln y_t \tag{3}$$

where ρ is the coefficient of autocorrelation and $\hat{\sigma}^2(\lambda,\mu,\theta,\rho)$ is the estimate of the residual variance conditional on the value of the autocorrelation and transformation parameters. The ML estimates of the parameters involved can be found by maximizing (3) by numerical methods or by searching (3) over the parameter space until a maximum of the likelihood function is reached. In the present study the latter method is employed.[3]

Estimation Results

Models A to E, as specified in the previous sections, are estimated using quarterly data covering the period from the first quarter of 1966 through the last quarter of 1981. The data used were obtained from *Livestock and Meat Situation* (U.S.D.A.) and from *Survey of Current Business* (U.S. Department of Com-

[2] This assumption is not strictly correct, since the power transformation employed does not permit negative values of the variables involved, which implies that the distribution of u_t is truncated, the truncation point being a function of the explanatory variables and of the transformation parameters. However, Draper and Cox have shown that symmetry in the distribution of the transformed variable is sufficient to guarantee robustness to nonnormality of the ML estimators.

[3] The four-dimensional (λ, μ, θ, ρ) grid search was carried out by an iterated generalized least square method similar to the IOLS method illustrated by Spitzer. The log-likelihood function was found to be steep at the first step and relatively flat at subsequent iterations, as may be gathered from the values reported in Table 2. On the implementation of grid search estimation using standard computer programs see Seaks and Layson.

TABLE 1. Maximum Likelihood Estimates of U.S. Beef Demand and Food Expenditure Equations.

Model	Dependent Variable (Transformed by λ)	Untransformed Explanatory Variables				Explanatory Variables Transformed by μ		
		Constant	D2	D3	D4	RPBI[a]	RPPK[a]	CPI
A	PCDBF	1.2245 (11.243)[b]	−0.53029 (0.60678)	2.2379 (0.67848)	0.58126 (0.60798)	−0.32016 (0.03367)	0.12513 (0.03528)	0.07445 (0.03626)
B	PCDBF	55.867 (3.2031)	−0.22128 (0.29046)	1.2108 (0.32857)	0.42210 (0.29071)	−0.02072 (0.00214)	0.00710 (0.00288)	
C	PCDBF	−126.89 (113.06)	−0.58083 (1.0658)	3.9949 (1.2046)	0.56783 (1.0697)	−0.01498 (0.00202)	0.00790 (0.00221)	
D	PCDBF	106.06 (14.386)	−0.13963 (0.46795)	2.0111 (0.52948)	0.42382 (0.45738)	−0.00792 (0.00115)	0.00351 (0.00128)	
E	PCEXF	−0.84877 (0.06082)	−0.00473 (0.00253)	−0.00738 (0.00291)	−0.00455 (0.00264)			

[a] These variables are deflated by CPI in model C and by CPIOF in model D.
[b] Values in parentheses are asymptotic standard errors.

merce). The maximum likelihood estimates are reported in Table 1. These estimates appear satisfactory in terms of sign, standard error, and fit.[4] The hypothesis of separability does not seem to add much to the analysis, as models B and D are statistically indistinguishable from A and C. The unrestricted models A and B, where homogeneity is not imposed, display a better fit in terms of the adjusted R^2 than models C and D, where the homogeneity condition is imposed.

The maximum likelihood method of estimation provides the ideal framework to test hypotheses regarding functional form and autocorrelation. It is well known that, if Ω is the unrestricted parameter space and ω is the restricted parameter space, the likelihood ratio testing procedure entails that the statistic

$$2[L(\hat{\Omega}) - L(\hat{\omega})] \qquad (4)$$

is asymptotically distributed as χ^2 with q degrees of freedom, where q is the number of restrictions imposed on ω (Theil, p. 396). Thus, the values of the alternative log-likelihood functions are reported in Table 2.

With regard to functional form, in the undeflated models A and B, the linear, log-log, and linear-log specifications are all rejected at the 0.01 probability level. When the homogeneity condition is imposed (equations C and D), the discrimination between functional form is less clear. Only the log-log formulation is rejected in both cases. In all four demand equations, the hypothesis of linearity in dependent and price variables and logarithmic form in the income term ($\lambda = \mu = 1$, $\theta = 0$) is the only one that is never rejected. Finally, the hypothesis that $\lambda = \mu = \theta$, that is that the Box-Cox transformation parameters can be reduced to a single parameter, is clearly rejected for models A and B, whereas it is accepted in models C, D, and E. As to the presence of autocorrelation, the hypothesis that $\rho = 0$ is rejected in all five models. This finding casts some doubts on previous applications of the Box-Cox transformation which have ignored the autocorrelation problem.

[4] The standard errors reported are those of the last GLS iteration, and therefore are conditional on the Box-Cox parameters. As Spitzer has shown, they are biased downward and strictly speaking cannot be used for hypothesis testing.

TABLE 1. Extended.

Explanatory Variables Transformed by μ			Explanatory Variables Transformed by θ		Transformation Parameters				Log Likeli-hood	$\bar{R}^2$
CPIOF	CPIF	CPINF	PCDY[a]	PCEXF[a]	λ	μ	θ	ρ		
			135.93 (22.854)		1.38	1.00	−1.25	0.49	28.968	0.873
−0.00224 (0.00326)				17.702 (3.3836)	1.16	1.38	−0.19	0.49	28.681	0.872
			615.75 (253.42)		1.56	1.97	−1.97	0.81	22.688	0.848
				70.460 (29.882)	1.31	1.94	0.94	0.87	21.105	0.841
	0.49694 (0.07739)	0.28983 (0.14232)	0.20835 (0.05091)		−0.06	−0.44	0.38	0.66	301.06	0.999

Parameter Stability Analysis

In order to facilitate the stability analysis, attention is concentrated on the stability of the parameter vector β. Consequently, it is assumed that parameters defining the functional form and the error structure are optimally estimated over the whole period, and thus λ, μ, θ and ρ are fixed at their estimated value in the stability analysis.[5]

Having transformed the variables by the estimated autocorrelation and Box-Cox coefficients, model (2) can be rewritten as

$$y = X\beta + e \qquad (5)$$

which in a more general varying parameter notation can be expressed as

$$y_t = x_t'\beta_t + e_t, \qquad t = 1, \ldots, T \qquad (6)$$

where y_t is the t^{th} observation of the dependent variable, x_t is a k-component vector of non-stochastic regressors, β_t is a k-component vector of unknown parameters, and e_t is the t^{th} component of the disturbances vector e, which, because of the autoregressive transformation, is now an independently normally distributed variate with variance σ^2.

For this form of the models A to E, the hypothesis that preferences for beef consumption did not change can be reduced to

$$H_0: \beta_1 = \beta_2 = \ldots = \beta_T = \beta$$

which can be subjected to a number of statistical tests. In particular, results of tests based on recursive residuals and on the F-distribution are presented below.

Recursive Analysis

Recursive residual analysis, introduced by Brown *et al.* and extended by Dufour (1982a), is particularly convenient as a starting point since it offers a flexible exploratory procedure sensitive to a variety of instability patterns. Recursive residuals are derived by estimating equation (5) recursively, that is by using the first K observations to get an initial estimate of β, and then gradually enlarging the sample,

[5] Here we extend to the functional form parameters one of Dufour's (1982a) suggestions, who recommended that ρ be estimated on the whole sample and fixed previous to the stability analysis on β. Although arbitrary, this procedure allows the entertained structural change to univocally affect the various demand elasticities, the analysis of which we are most interested in.

TABLE 2. Log-Likelihood Values under Parameter Restrictions.

Model	Unrestricted Log-Likelihood	Restricted Log-Likelihood Functional Form Tests					Autocorrelation Test
	$L(\hat{\lambda},\hat{\mu},\hat{\theta},\hat{\rho})$	$L(\lambda = \mu = \theta = 1,\hat{\rho})$	$L(\lambda = 1, \mu = \theta = 0,\hat{\rho})$	$L(\lambda = \mu = \theta = 0,\hat{\rho})$	$L(\lambda = \mu = 1; \theta = 0,\hat{\rho})$	$L(\lambda = \mu = \theta,\hat{\rho})$	$L(\hat{\lambda}, \hat{\mu}, \hat{\theta}, \rho = 0)$
A	28.968	20.903**	19.069**	17.676**	27.573	21.226**	21.596**
B	28.681	21.328**	17.743**	16.403**	28.133	22.062**	21.183**
C	22.688	20.157	18.740*	17.131*	20.699	22.003	0.67043**
D	21.105	19.927	18.257	16.892*	19.902	20.912	−7.0259**
E	301.06	295.53*	284.34**	300.141	290.94**	300.47	284.72**

* The restrictions imposed on the parameters are rejected at the 0.05 level based on the χ^2 distribution.
** The restrictions imposed on the parameters are rejected at the 0.01 level based on the χ^2 distribution.

adding one observation at a time and reestimating β at each step.[6] In this way it is possible to get (T − K) estimates of the vector β, and correspondingly (T − K − 1) forecast errors of the type

$$v_r = y_r - x_r'b_{r-1}, \qquad r = K + 1, \ldots, T \qquad (7)$$

where b_{r-1} is an estimate of β based on the first r − 1 observations. Since it can be shown that these forecast errors under H_0 have mean zero and variance $\sigma^2 d_r^2$, where d_r is a scalar function of the explanatory variables, the quantity

$$w_r = v_r/d_r, \qquad r = K + 1, \ldots, T \qquad (8)$$

gives a set of standardized prediction errors, which were called "recursive residuals" by Brown *et al.*, and which under H_0 are identically independently distributed normal variates with mean zero and variance σ^2. Given that parameter instability would perturb the stochastic properties of recursive residuals, Dufour (1982a) suggests testing for departures from this null distribution.

Table 3 summarizes the results of tests performed on one-step ahead recursive residuals obtained by forward recursion.[7] The t-test is performed to test that the mean of the recursive residuals is not different from zero, their expected value under H_0. The nonparametric Sign and Wilcoxon tests are used to test for the

[6] Recursive estimation, as pointed out by Dufour (1982a), can be viewed as a special case of Kalman filtering, a technique already employed by Chavas to study structural change in the U.S. demand for meat.

[7] The recursive residuals obtained by backward recursion were also tested. Since the results were similar to those based on the forward recursion, they are not reported.

TABLE 3. Non-Parametric Tests on Recursive Residuals.

Model	t-Test	Sign Test Value	Sign Test Z-Value	Wilcoxon Test Value	Wilcoxon Test Z-Value	Runs Test Value	Runs Test Z-Value
A	−1.083	24	−1.069	649	−1.215	29	0.135
B	−0.011	24	−1.069	773	−0.204	29	0.135
C	−0.692	28	−0.132	754	−0.576	32	0.802
D	0.234	26	−0.662	851	0.195	30	0.267
E	0.526	28	−0.132	880	0.425	24	−1.336

Note: Number of recursive residuals is 56 for models A and B, and 57 for models C, D, and E.

symmetry and zero median implied by the normality of the recursive residuals, while the Runs test is aimed at testing the independence of the distribution of recursive residuals (Dufour, 1981).[8] From Table 3, it is evident that for all five models the distribution of recursive residuals implied by the constancy of parameters hypothesis cannot be rejected.

Recursive residuals can also be used to construct other potentially useful tests. In particular, Brown *et al.* recommended the use of two graphs based on recursive residuals, CUSUM and CUSUMSQ.

CUSUM involves the plot of the quantity

$$W_r = \frac{1}{\hat{\sigma}} \sum_{t=K+1}^{r} w_t, \quad r = K + 1, \ldots, T \qquad (9)$$

where $\hat{\sigma}$ is the estimated standard deviation based on the full sample.

CUSUMSQ is based on the square of the recursive residuals, and involves the plot of the quantity

$$Sr = \sum_{t=K+1}^{r} w_t^2 \Big/ \sum_{t=K+1}^{T} w_t^2, \quad r = K + 1, \ldots, T. \qquad (10)$$

Since CUSUM retains the information of the sign of w_t, it is aimed at detecting systematic movements of the vector β_t. CUSUMSQ, on the other hand, is designed to detect haphazard movements in the coefficient vector.[9] Under the null hypothesis of no parameter change, probabilistic bounds for both W_r and S_r can be determined, so that H_0 is rejected if the plot crosses the boundaries associated with the chosen probability level.

CUSUM and CUSUMSQ plots are not presented but, as reported by Moschini and Meilke, in no case do these plots cross the probability boundaries at the 0.05 level, nor are the plots indicative of any particular point of discontinuity in the coefficients. Overall, the recursive residual analysis suggests that the hypothesis of constancy of the coefficient vector β cannot be rejected.

Tests Based on the F-Distribution

By making more precise assumptions about the pattern of parameter instability, it is possible to use tests based on the F-distribution, such as the Chow test and the Farley-Hinich test.

The Chow test is probably the best known test for detecting structural change.[10] It amounts to testing whether a regression equation is the same between two disjoint subperiods. Thus the test involves breaking the sample at the point at which the parameter shift is believed to have taken place, and then testing for equality between the coefficient vectors in the two samples using an F-test. The test need not include all of the coefficients, so the hypothesis of structural change can be formulated to affect only a subset of the vector β.

Since the point of structural change is typically unknown, Farley and Hinich (F-H) suggest approximating the discrete shift in the parameters at an unknown point by a linear continuous shift; that is, by letting $\beta_t = \beta + t\delta$. Equation (6) then becomes

$$y_t = x_t'(\beta + t\delta) + e_t \qquad (11)$$

and the F-H test reduces to a test of the null hypothesis H_0: $\delta = 0$ against H_1: $\delta \neq$

[8] Although these non-parametric tests have a known exact distribution, for sample sizes like ours the normal approximation works well (Hájek). For simplicity, then, in Table 4 besides the value of the tests the corresponding standardized z-values are reported.

[9] Based on CUSUM and CUSUMSQ, Hassan and Johnson (1979b) concluded that quarterly meat demand in Canada was relatively stable.

[10] See Chow and, for a highly concise and effective presentation, Fisher. A generalization of this test to several regressions with arbitrary sample size is found in Dufour (1982b).

TABLE 4. Farley-Hinich and Chow Tests for Structural Change.

		F-Statistics			Critical Value	
Model	Coefficient Subset Tested	Farley-Hinich	Chow 1966–73 1974–81	Chow 1966–78 1979–81	Degrees of Freedom	$F^a_{0.05}$
A	All coefficients	1.43	1.66	1.19	(8,48)	2.14
	Constant and dummies	0.88	1.08	0.45	(4,52)	2.55
	Slope coefficients	1.53	2.27	0.91	(4,52)	2.55
	Price coefficients	2.07	3.00*	0.18	(3,53)	2.78
B	All coefficients	1.25	2.13	0.41	(8,48)	2.14
	Constant and dummies	1.05	0.69	0.45	(4,52)	2.55
	Slope coefficients	1.38	2.76*	0.11	(4,52)	2.55
	Price coefficients	1.40	3.73*	0.14	(3,53)	2.78
C	All coefficients	2.47*	1.94	1.82	(7,50)	2.20
	Constant and dummies	1.35	0.94	2.13	(4,53)	2.55
	Slope coefficients	3.05*	2.72	2.28	(3,54)	2.78
	Price coefficients	4.41*	3.65*	3.31*	(2,55)	3.16
D	All coefficients	1.92	1.82	0.95	(7,50)	2.20
	Constant and dummies	1.16	0.98	1.40	(4,53)	2.55
	Slope coefficients	2.70	2.74	1.79	(3,54)	2.78
	Price coefficients	4.12*	4.17*	2.70	(2,55)	3.16
E	All coefficients	0.64	0.50	0.71	(7,50)	2.20
	Constant and dummies	0.29	0.79	0.32	(4,53)	2.55
	Slope coefficients	0.53	0.76	0.59	(3,54)	2.78
	Price coefficients	0.80	1.12	0.67	(2,55)	3.16

[a] Computed by interpolation using the reciprocal of the degrees of freedom.
* The calculated statistic is greater than the critical value at the 0.05 probability level.

0.[11] Again, the test may involve the full set or only a subset of the coefficients β.

Table 4 reports the computed F-H and Chow tests. Both the F-H and Chow tests have been performed on four different sets of coefficients for all five models: a) the full set of coefficients; b) the constant and seasonal dummies; c) the price and income coefficients; and d) only the price coefficients. The Chow test has been performed both by breaking the sample in the middle (at the end of 1973) and by isolating the last three years.

When the full set of coefficients are tested, only the F-H test for model C indicates structural change. Constants and dummies taken alone are never significant. The hypothesis that only price and income coefficients are subject to structural change is accepted by the F-H test in model C and by the Chow test (with break at 1973) for model B. Finally, the hypothesis that the structural change affected only the price coefficients is accepted for all four demand equations by the Chow test (break at 1973), and for models C and D using the F-H test. The broad consistency of the results for the four different specifications of demand suggests that the price coefficients of demand have actually been following two separate regimes. Moreover, when the income coefficient is tested along with the prices, the F-test is insignificant in most cases. This indicates that the income coefficient has

[11] Farley *et al.* provide Monte Carlo evidence that their test is robust with respect to gradual parameter shifts in one or more parameters. Moreover, they show that the Chow test performed by breaking the sample in the middle dominates the F-H test if the (unknown) structural change point is approximately within the middle 20 percent of the record, whereas the converse holds if the shift point moves to either end of the record.

TABLE 5. Estimated Elasticities of U.S. Beef Demand under Constancy of Parameters and under Structural Change—Selected Quarters.

		RPBF		RPPK		CPI/CPIOF[a]		CPIF	CPINF	PCDY/PCEXF[b]	
Model	Evaluation Point	Constant Parameters	Structural Change[c]	Constant Parameters	Structural Change[c]	Constant Parameters	Structural Change[c]	Constant Parameters	Constant Parameters	Constant Parameters	Structural Change[c]
A	1966 1	−0.29	−0.44	0.11	0.06	0.08	0.28			0.47	0.49
	1971 1	−0.33	−0.48	0.09	0.05	0.09	0.32			0.29	0.30
	1976 1	−0.38	−0.35	0.15	0.19	0.10	0.04			0.14	0.15
	1981 1	−0.84	−0.77	0.21	0.27	0.22	0.08			0.10	0.11
	mean	−0.45		0.13		0.12				0.19	
B	1966 1	−0.21	−0.31	0.07	−0.07	−0.03	0.23			0.45	0.48
	1971 1	−0.25	−0.37	0.05	−0.06	−0.03	0.27			0.40	0.43
	1976 1	−0.35	−0.34	0.12	0.19	−0.05	−0.10			0.32	0.34
	1981 1	−0.89	−0.87	0.16	0.26	−0.12	−0.23			0.37	0.39
	mean	−0.39		0.09		−0.05				0.37	
C	1966 1	−0.60	−0.37	0.29	0.00	−0.24	−0.15			0.55	0.52
	1971 1	−0.51	−0.31	0.13	0.00	−0.04	−0.08			0.42	0.39
	1976 1	−0.42	−0.46	0.22	0.32	−0.08	−0.12			0.28	0.26
	1981 1	−0.66	−0.72	0.14	0.20	0.15	0.17			0.37	0.35
	mean	−0.56		0.18		0.02				0.36	
D	1966 1	−0.59	−0.37	0.24	−0.06	−0.20	−0.01			0.55	0.44
	1971 1	−0.58	−0.36	0.12	−0.03	−0.08	−0.04			0.54	0.43
	1976 1	−0.41	−0.47	0.18	0.30	−0.21	−0.18			0.44	0.35
	1981 1	−0.65	−0.75	0.12	0.19	−0.07	0.08			0.60	0.48
	mean	−0.57		0.15		−0.10				0.52	
E	1966 1							0.48	0.29	0.36	
	1971 1							0.45	0.26	0.41	
	1976 1							0.38	0.23	0.49	
	1981 1							0.33	0.20	0.57	
	mean							0.40	0.24	0.38	

[a] CPI is used in models A and C, CPIOF in models B and D. For models C and D the elasticities are obtained from the homogeneity condition.
[b] PCDY is used in models A, C, and E, PCEXF in models B and D.
[c] Based on parameters estimated with Chow regressions with break at the end of 1973 for the price coefficients.

been relatively stable. Also, comparison of the Chow test, with a break at 1973(4) and with a break at 1978(4), suggests that if there was a structural change in the demand for beef, it occurred early rather than late in the 1970s.

It is interesting to look at the effects of the entertained structural change on estimated elasticities. Thus in Table 5 price and income elasticities, evaluated for the first quarter of selected years and at the means of the variables involved, are reported. These were estimated both under the constancy of parameters assumption and under the assumption that a structural change, affecting the price coefficients only, occurred in 1973(4). The elasticities at the mean point are broadly consistent with previous studies (George and King; Pope *et al.*; Blaylock and Smallwood). What is perhaps more interesting is the behavior of the elasticities over time.[12]

[12] For our model, income elasticity falling as income rises requires $(\theta - \lambda\epsilon_y) < 0$, whereas own price elasticity rising in absolute value as the own price rises requires $(\mu - \lambda\epsilon_p) > 0$. Given the sign of the parameters estimated, it is easily verified that these conditions, which are usually held as normal (Gemmill), are satisfied.

Consider first the case of constant parameters. In the unrestricted models A and B, the own price elasticity shows a net increase in absolute value, especially in the last five years. The variability of the own price elasticity is much smaller in the restricted models C and D, although there is still a 50 percent increase in its value in the last five years. The cross price elasticities in the case of constant parameters are less consistent across the different models. The elasticities with respect to pork price increases through the sample period in models A and B, indicating an increasing substitutability between beef and pork, whereas, it decreases in the restricted models C and D. The elasticities with respect to CPI and CPIOF are even less clear, switching from substitutability to complementarity in some cases. The income elasticity shows a sizable decrease in model A, whereas, it is fairly stable for models B, C, and D. As for the first budgeting stage of model E, the positive sign of the own-price elasticity of food expenditure implies that the own-price elasticity of the aggregate commodity "food" is in the inelastic region.

The hypothesis of structural change does not affect the own-price elasticity of beef demand, which still displays a strong increase in absolute value in the last few years. What does change is the cross elasticity with respect to pork price, and indeed the results of the four models are in this case consistent. Under structural change this elasticity is close to zero in the first sub-sample, whereas it shows a higher degree of substitutability in the second part of the sample.

Overall then, it would seem that consumers have moved to a point on their demand surface characterized by higher own elasticity and higher cross price elasticity with respect to other meat. These results are partly in contrast with those obtained by Chavas who, relative to the demand for beef, found strong evidence of structural change, a decreasing (in absolute value) own-price elasticity, increasing cross-price elasticities, and a dramatically decreasing income elasticity. These differences in results can be accounted for by the different methodology used, the different sample period, and the fact that Chavas's base model is a constant elasticity one so that his evidence of structural change may partly reflect a functional form problem.

Conclusions

Providing a definite answer to the question of whether there has been a change in the structure of consumer preferences affecting U.S. beef consumption is difficult, especially because this hypothesis can be confused with many types of model misspecification. The results of our analysis do not support any strong conclusions, although the following observations seem justified.

First, the extended Box-Cox model shows that many functional forms introduce some degree of misspecification. Moreover, the hypothesis of no serial correlation is rejected for all the models analyzed. This suggests the inaccuracy of many applications of the Box-Cox transformation which have neglected the autocorrelation problem.

Second, with regard to the question of structural change, the analysis based on recursive residuals gives no particular indication of parameter instability. The tests based on the F-distribution, on the other hand, provide some evidence that structural change may have occurred. This parameter change, however, appears to have taken place early rather than late in the 1970s, to have been confined mainly to the price coefficients, and to affect primarily the cross-price elasticities. Overall, however, the evidence of structural change is weak, and suggests that the recent decline in beef consumption may be attributed to changed market conditions and is thus of a reversible nature.

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