Structural Change in U.S. Meat Demand: Further Evidence

by

Giancarlo Moschini and Karl D. Meilke

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Consumption patterns for meat products have changed considerably over the last two decades. This development is illustrated in Figure 1, which shows quarterly per capita meat consumption levels (pounds of retail weight for beef, pork, and chicken, and expenditure in 1987 dollars for fish and seafood). The most striking feature of this picture is the steady increase in chicken consumption, which has virtually doubled in the last twenty years. Also noticeable is the sharp decline in beef consumption that has taken place in the second half of the 1970s. While the tremendous gain in productivity experienced by the poultry sector, which has translated in lower real retail prices for poultry, is an obvious explanation for this phenomenon, there has been a growing suspicion that other forces are also at work. Specifically, it is hypothesized that the changing meat consumption patterns reflect, at least partly, an evolution in the underlying consumers' preferences, driven by a growing awareness of the health hazard of large intakes of cholesterol and other saturated fats.

This hypothesis carries considerable interest for the red meat industry, as it implies a more unfavorable climate warranting adjustments in both production and marketing strategies. Also, changing preferences add problems to the difficult task of forecasting meat prices through structural models, as changes of unknown form in the data generating process invalidate commonly used econometric techniques. Finally, as the issue hits the popular press, less distinct and yet worrisome fears seem to emerge. As Smith (1987) notes, "Cowboys, not chicken farmers tamed the west." Will this country still be the same as it shifts from a nation of red meat eating "cowboys" to a nation of white meat eating "yuppies"?

Considerable research has been devoted to analyzing the hypothesis of structural change in meat demand, including Braschler (1982), Chalfant and Alston (1987), Chavas (1983), Dahlgran (1987), Moschini and Meilke (1984), and Nyamkori and Miller (1982), Thurman (1987), and Wohlgenant (1985). The evidence from these studies is mixed, a not surprising outcome given the variety of methods and data employed.[1] Meanwhile, it is becoming clear that determining whether consumers' preferences have indeed changed is an elusive task (Chavas, 1986). Since important features of the model generating the data are not known (e.g. the shape of utility functions) or are deliberately simplified for empirical tractability (e.g. by aggregating across commodities and/or across consumers), there is always the possibility that evidence of structural change may reflect model misspecification of some kind. Yet, given the scope of empirical analysis, the issue of structural change still appears a relevant one, although possibly limited to the properties of an estimated model. At a minimum, evidence of the existence and nature of structural change will improve the quality of inferences from such models.

Given the above, the objective of this paper is to provide further evidence of the nature of structural change in meat demand. As in Christensen

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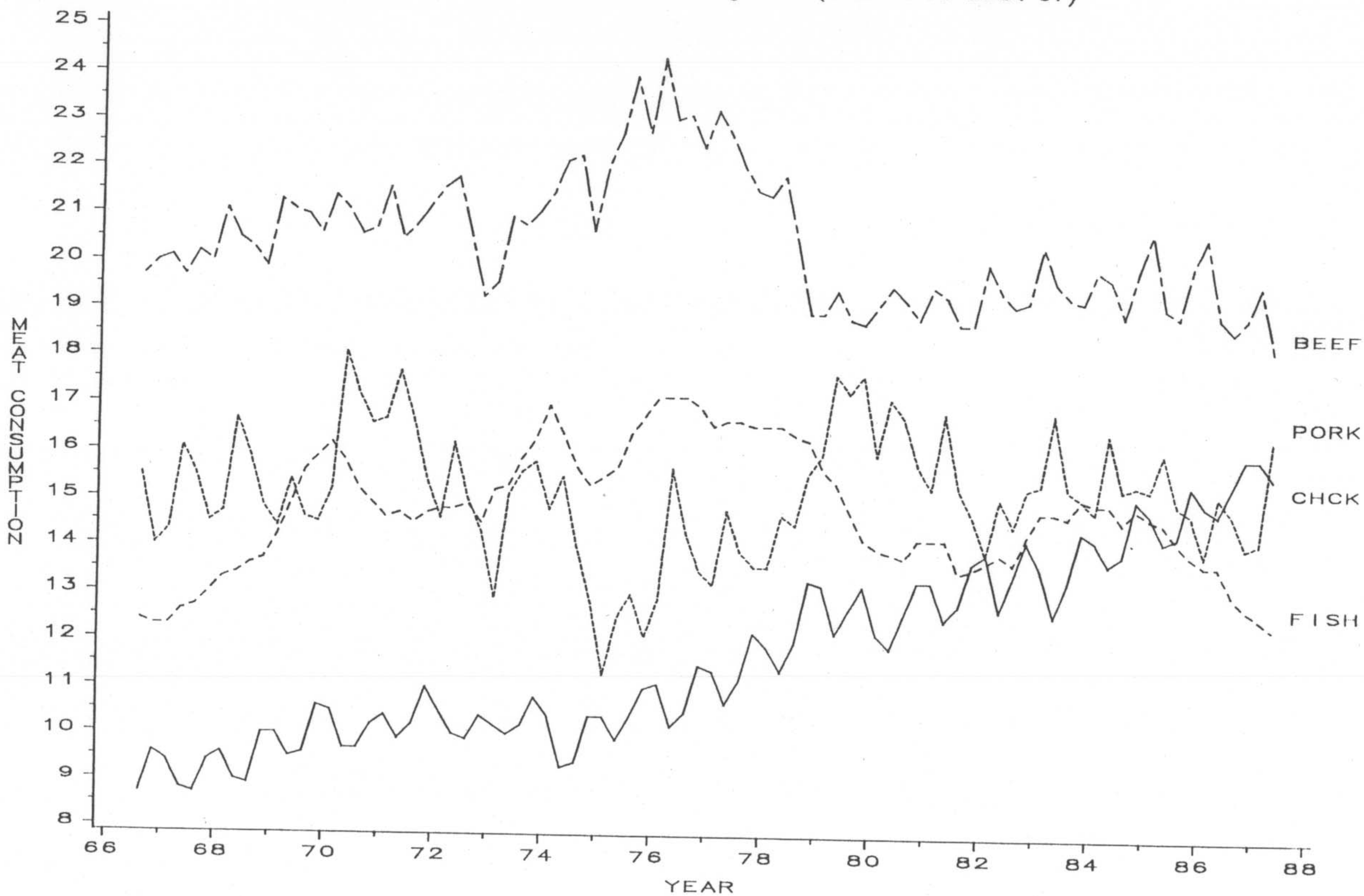
Figure 1 – U.S. Meat Consumption (1967 I to 1987 IV)



and Manser (1977), attention is restricted to the substitutability amongst four meats (beef and veal, pork, chicken, and fish) by assuming weak separability between these products and all other goods. The analysis is carried out within a model satisfying acceptable definitions of flexibility and the restrictions of an underlying utility maximization process. The data are quarterly, and spans the period 1967 to 1987. The focus of the analysis is to test for structural change in a model that can provide information on the timing and time path of this change, and to investigate the bias of structural change on consumption patterns and estímated elasticities.

Model Specification

The combined development of duality theory and flexible functional forms allows a number of system approaches to modeling consumer behavior (Deaton, 1986; Johnson, Hassan, and Green, 1984). In this paper we employ the linear version of the Almost Ideal Demand System (AIDS) derived by Deaton and Muellbauer (1980). This model provides a first-order approximation to an arbitrary demand system, and satisfies perfect aggregation conditions over consumers. In addition, the linearity of the AIDS model is an attractive feature for the econometric analysis of structural change addressed in this paper. The stochastic version of the system of share equations generated by the AIDS model can be written as:

$$w_{it} = \alpha_i + \Sigma_j \beta_{ij} p_{jt} + \beta_i x_t + \Sigma_k \alpha_{ik} D_k + e_{it} \tag{1}$$

where (i,j) index the n goods (4 in our case), t indexes time, $p \equiv \ln(P)$ with P denoting nominal prices, the i^{th} good share satisfies $w_i \equiv (P_i Q_i/X)$ with Q_i being the quantity demanded and X total expenditures on the four goods, $x \equiv \ln(X/P^*)$ with P^* being a price index which is approximated by Stone geometric index, k indexes three seasonal dummies denoted by D, and e is an error term. The theoretical properties of homogeneity in prices and income, and symmetry of the cross effects of demand functions, imply the following parametric restrictions on (1):

$$\Sigma_j \beta_{ij} = 0 \tag{2.1}$$

$$\Sigma_i \alpha_i = 1, \qquad \Sigma_i \beta_{ij} = 0 , \qquad \Sigma_i \beta_i = 0 , \qquad \Sigma_i \alpha_{ik} = 0 \tag{2.2}$$

$$\beta_{ij} = \beta_{ji} \tag{2.3}$$

where (2.1) follows from the zero homogeneity of the share equations, (2.2) is the adding-up condition (implied by homogeneity through Euler's theorem), and (2.3) is the Slutsky symmetry condition.[2]

Under the hypothesis of no structural change, and assuming that the AIDS model provides a satisfactory approximation of the true demand system, the set of parameters $\{\alpha,\beta\}$ in equation (1) gives a full representation of the underlying utility maximization process. It follows that changing preferences can be characterized by allowing this set of parameters to change over time. Without specific knowledge concerning the way parameters change over time, as a first approximation it can be assumed that the time path of the parameters is the same. If we denote this path as h_t, equation (1) can be reparameterized as:[3]

$$(3) \qquad w_{it} = \alpha_i + \gamma_i h_t + \Sigma_j (\beta_{ij} + \delta_{ij} h_t) p_{jt} + (\beta_i + \delta_i h_t) x_t + \Sigma_k (\alpha_{ik} + \gamma_{ik} h_t) D_k + e_{it}$$

Thus, a test for parameter constancy (absence of structural change) can be performed by testing the restrictions $\gamma = 0$ and $\delta = 0$. If the time path h_t is approximated by a vector whose elements are 0 for the first half of the sample and 1 for the latter half, one obtains the Chow (1960) test for structural change. If the unknown path is approximated by the linear path t/T, one obtains the test studied by Farley and Hinich (1970). To shed some light on the actual shape of this path, we follow Ohtani and Katayama (1986) and postulate that the structural change path can be approximated by:

(4.1) $\qquad h_t = 0 \qquad$ for $t = 1,\ldots,\tau_1$

(4.2) $\qquad h_t = (t-\tau_1)/(\tau_2-\tau_1) \qquad$ for $t = \tau_1+1,\ldots,\tau_2-1$

(4.3) $\qquad h_t = 1 \qquad$ for $t = \tau_2,\ldots,T$

It is clear that (4) implies that (3) becomes a gradual switching regression model where τ_1 represents the end-point of the first regime and τ_2 is the starting point of the second regime, and where the transition path between the two regimes is linear. To identify the structural change path, τ_1 and τ_2 are viewed as parameters to be estimated. Note that if $\tau_2=\tau_1+1$, the change in the coefficient vector is abrupt, while $\tau_1=0$ and $\tau_2=T$ produces the Farley-Hinich parameter path.[4]

Finally, for the purpose of estimation the set of equations (3) is expressed in first difference form. The use of differencing in econometric analysis of time series has been advocated since Box and Jenkins (1970) to avoid the "spurious" regression problem. In the present case this approach can be defended as a reasonable choice since it allows a very parsimonious representation, in the spirit of Hendry and Mizon (1978), of the dynamic behavior to be expected in frequent time series data. Thus, the model that is estimated is:

$$(5) \qquad \Delta w_{it} = \gamma_i \Delta h_t + \Sigma_j [\beta_{ij} \Delta p_{jt} + \delta_{ij} \Delta(h_t p_{jt})] + \beta_i \Delta x_t + \delta_i \Delta(h_t x_t) + \Sigma_k [\alpha_{ik} \Delta D_k + \gamma_{ik} \Delta(h_t D_k)] + u_{it}$$

with the parameter path h_t given by equations (4). Because of the adding-up condition $\Sigma_i \Delta w_i = 0$, only n-1 of these equations can be used for estimation. Under the assumption that the error terms u_{it} in (5) are multinormally distributed but contemporaneously correlated with:

(6.1) $\qquad E(u_{it}) = 0$

(6.2) $\qquad E(u_{it} u_{jt}) = \omega_{ij}$

(6.3) $\qquad E(u_{it} u_{js}) = 0 \qquad$ for $t \neq s$

maximum likelihood estimation can be performed for any given parameter path h_t. Given that h_t in (6) has discontinuous derivatives with respect to τ_1 and τ_2, estimates of these two parameters are obtained by searching the likelihood function over the range of interest of (τ_1,τ_2). Under the above stochastic assumptions, the obtained maximum likelihood estimator is consistent, asymptotically normal, and asymptotically efficient, and is independent of which equation was deleted (Barten, 1969).

Data

The data utilized are quarterly disappearances and retail prices. For beef and pork, quantities are per capita disappearance in retail weight as published by the U.S. Department of Agriculture in Livestock and Meat Situation (LMS) and in Livestock and Poultry Situation (LPS). The quantity of chicken is the total of young and mature chicken per capita disappearance as published by the U.S. Department of Agriculture in Poultry and Eggs Situation (PES) and in LPS. For fish, personal consumption expenditure for fish and seafood on a quarterly basis was obtained from unpublished U.S. Department of Commerce data, and this was expressed on per capita terms using estimates of U.S. population published on the Survey of Current Business by the U.S. Department of Commerce. Retail prices for beef (choice) and pork as published in LMS and LPS, the retail price of frying chicken published in PES and LPS, and the CPI for fish and seafood from the U.S. Department of Commerce were used. For the purpose of estimation, prices and real income were normalized by their sample mean previous to the logarithmic transformation, and the seasonal dummy variables are specified for the first three quarters.

Results

Using the data described above for the period spanning the first quarter of 1967 to the last quarter of 1987, the first difference AIDS model in (5) was estimated using the maximum likelihood procedure available in SHAZAM 6.0. To estimate the parameters of the structural change path, the system of equations was estimated for all the combinations of τ_1 from 1971(1) to 1983(4) and τ_2 from 1971(2) to 1987(4) (for $\tau_1 < \tau_2$). This set of combinations ensures that all the parameters in (5) are identified, and requires estimating the system of equations a total of 2,158 times. The estimated coefficients for the three-equation system, conditional on the estimated τ parameters, are reported in Table 1, along with some single equation statistics. The fit of the model is good for the pork and chicken equations, and only satisfactory for the beef. In any case, one should bear in mind that these R^2 coefficients measure the fit to a left-hand-side defined as a first difference. The Durbin-Watson (DW) statistics reported in Table 1 can be interpreted as a rough test of the validity of the dynamic specification, and of the specification of structural change. The test does not reject the hypothesis of no autocorrelation for the three equations, suggesting that the specification of the model is acceptable. In particular, the use of first differences appears a useful simplified dynamic specification device, though only one of numerous dynamic specifications that could in principle be used. Not much can be said about the individual coefficient estimates reported in Table 1, and more can be understood by doing structural tests and computing relevant elasticities.

Concerning the issue of structural change, the value of the parameters defining the path of structural change that maximize the set of likelihood functions are:

Table 1 - Maximum likelihood parameter estimates for the time varying AIDS model

Equation	intercept	deflated price: beef	deflated price: pork	deflated price: chicken	expenditure	seasonal dummies: 1st qrt	seasonal dummies: 2nd qrt	seasonal dummies: 3rd qrt	R^2	DW
Beef	n.a.	0.07189 (0.02374)	0.02982 (0.01975)	-0.05486 (0.00850)	0.11727 (0.04424)	0.00842 (0.00172)	0.01412 (0.00259)	0.01728 (0.00190)		
	-0.02980 (0.01348)	0.00351 (0.03134)	-0.01312 (0.02594)	0.01096 (0.01159)	0.08197 (0.06539)	0.00234 (0.00261)	0.00001 (0.00337)	0.00053 (0.00239)	0.76	1.87
Pork	n.a.	0.02982 (0.01975)	-0.00105 (0.02053)	-0.01180 (0.00793)	0.01121 (0.04036)	-0.00919 (0.00158)	-0.02409 (0.00240)	-0.02596 (0.00173)		
	0.00883 (0.01254)	-0.01312 (0.02594)	0.03459 (0.02695)	-0.01881 (0.01108)	-0.05260 (0.05956)	-0.00465 (0.00239)	-0.00013 (0.00311)	0.00165 (0.00217)	0.90	2.19
Chicken	n.a.	-0.05486 (0.00850)	-0.01180 (0.00793)	0.08082 (0.00647)	-0.07359 (0.01929)	0.00065 (0.00066)	0.00938 (0.00100)	0.00767 (0.00073)		
	0.00869 (0.00513)	0.01096 (0.01159)	-0.01881 (0.01108)	0.00505 (0.00984)	-0.00980 (0.02696)	0.00260 (0.00100)	-0.00011 (0.00129)	-0.00082 (0.00092)	0.94	1.92

Note: For each equation, the first row of coefficients refer to the fixed component of the parameter, while the second row is the time varying component of the parameters;
the asymptotic standard errors (conditional on the estimated value of time path parameters) are reported in parentheses;
maximized log-likelihood = 1,125.28 .

$$\hat{\tau}_1\text{=1975 IV and }\hat{\tau}_2\text{=1976 III.}$$

This result suggests a rather abrupt change of regimes roughly in the middle of the observation period, with the second regime starting at the peak of beef consumption (per capita beef consumption reached a maximum 24.4 lb/quarter in 1976 III before experiencing a sharp 20 percent decline in less than three years). To investigate the significance of the structural change, Table 2 reports likelihood ratio tests for the hypothesis of constancy of the parameter vector over time. The hypothesis of no structural change is rejected at the 0.05 significance level, and this rejection is more pronounced for the seasonal parameters than for the structural parameters of the AIDS model. Thus, the evidence is that a constant set of parameters cannot be postulated to rationalize consumer behavior within the assumed model, which suggests some degree of structural change occurring over the period considered.

The search procedure implemented to estimate the parameters of the structural change path does not yield standard errors for the estimated τ parameters. However, it is possible to construct a confidence region for these two parameters over the explored parameter space. If we let (τ_1^o,τ_2^o) denote any pair of structural change parameters, (τ_1^o,τ_2^o) belong to the 95 percent confidence region if:

$$2[L(\hat{\tau}_1,\hat{\tau}_2)-L(\tau_1^o,\tau_2^o)] \leq \chi^2_{0.05}(2),$$

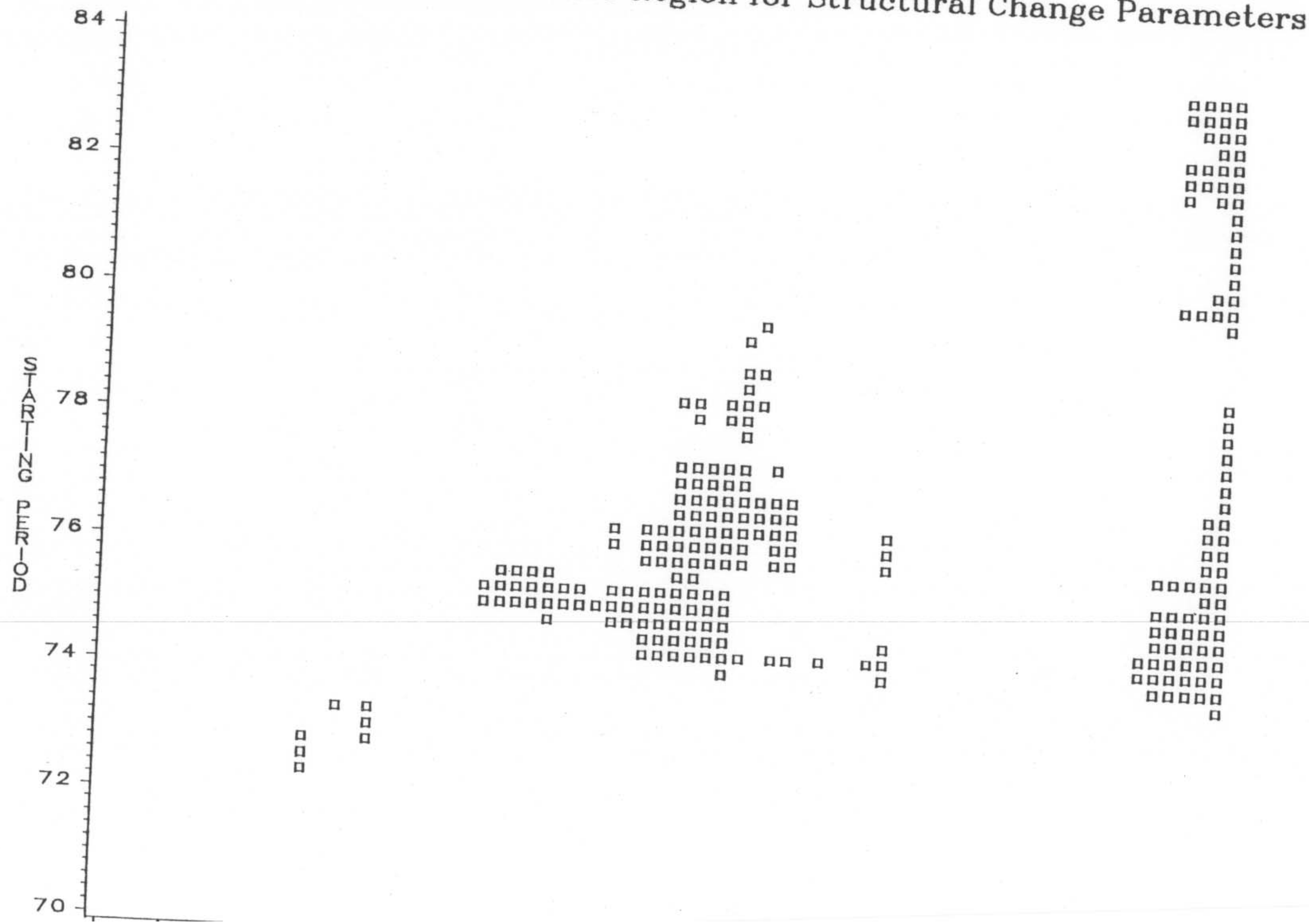
where L(.) denotes the log-likelihood function. Of the 2,158 combinations of (τ_1,τ_2) considered, 240 fall in this confidence region, and they are illustrated in Figure 2. While the estimated τ_1 and τ_2 suggest a sharp transition between two regimes, it is clear from Figure 2 that a variety of structural change paths cannot be rejected. In addition to a break point in the middle of the sample, Figure 2 suggests two other likely paths, one starting in the mid-1970's and the other starting in the early 1980's, both extending to the end of the period considered.

Given the evidence for structural change, it is interesting to see whether this significantly biases consumption patterns and the elasticities of demand. Concerning consumption patterns, and in analogy to measures of structural change in production (Binswanger, 1974), one can attribute bias to structural change if it significantly changes expenditure shares. If we let

Table 2 - Likelihood Ratios for Absence of Structural Change

Hypothesis	Number of Restrictions	Likelihood Ratio	$\chi^2_{0.05}$
No structural change	18	45.06	28.87
No structural change in structural parameters	9	18.46	16.92
No structural change in seasonal parameters	9	24.72	16.92

Figure 2 – 95 Percent Confidence Region for Structural Change Parameters



w_i^b denote the i^{th} good share before structural change, and w_i^a the same share after structural change, an obvious measure of bias is $B_i \equiv w_i^a - w_i^b$. Structural change will be biased against the i^{th} good if it resulted in a lower demand for this good, which would be implied by $B_i < 0$. Conversely, $B_i > 0$ implies that structural change is biased in favor of the i^{th} good. Evaluating the shares at the sample mean of prices and expenditures (such that p_{jt}=0 and x_t=0), this measure reduces to:

$$(7) \qquad B_i = \gamma_i + \Sigma_k \gamma_{ik} D_k$$

Table 3 reports this estimated bias, and the associated standard errors, for each quarter and for an average of the four quarters. Structural change appears to be significantly biased against beef, and in favour of chicken and fish, while it is neutral for pork. The hypothesis that the measures of bias reported in Table 3 are simultaneously different than zero can be tested by the Wald test (Engle). This test, distributed as χ^2 with 3 degrees of freedom because of the adding-up property, shows that the hypothesis of no bias of structural change can be rejected at the 0.05 significance level.

Finally, a problem of interest concerns the question of whether structural change has significantly affected demand elasticities. Marshallian elasticities that reflect the effects of structural change are computed as:

$$(8.1) \qquad \epsilon_{ii} = (\beta_{ii}+\delta_{ii})/w_i^a - (\beta_i+\delta_i) - 1$$

$$(8.2) \qquad \epsilon_{ij} = (\beta_{ij}+\delta_{ij})/w_i^a - (\beta_i+\delta_i)(w_j^a/w_i^a)$$

$$(8.3) \qquad \epsilon_{ix} = (\beta_i+\delta_i)/w_i^a + 1$$

Elasticities that reflect demand response before structural change are obtained from (8) by setting the δ parameters to zero and substituting w_i^b for

Table 3 - Biases of Structural Change

Share	Bias: 1st quarter	2nd quarter	3rd quarter	4th quarter	average
Beef	-0.0275	-0.0298	-0.0293	-0.0298	-0.0291
	(0.0130)	(0.0125)	(0.0133)	(0.0135)	(0.0130)
Pork	0.0042	0.0087	0.0105	0.0088	0.0080
	(0.0120)	(0.0125)	(0.0124)	(0.0125)	(0.0120)
Chicken	0.0113	0.0086	0.0079	0.0087	0.0091
	(0.0049)	(0.0048)	(0.0051)	(0.0055)	(0.0049)
Fish	0.0199	0.0125	0.0109	0.0123	0.0119
	(0.0052)	(0.0049)	(0.0054)	(0.0055)	(0.0052)
Wald Test	12.61	12.45	9.02	10.45	11.11

Note: Standard errors are reported in parentheses.

w_i^a. The mean of the estimated share over the whole period is used to represent w_i^a, while w_i^b is obtained by subtracting from w_i^a the average bias measure reported in Table 3. The elasticities obtained are reported in Table 4, along with standard errors computed as in Chalfant (1987) by treating expenditure shares as constants. The demand for beef and pork appear much more elastic than chicken and fish. Notably, beef is the only superior good in both regimes. The cross-price elasticities show more complementarity relationships than expected, with 10 of the 12 cross-price elasticities taking a negative sign in both regimes. It should be noted, however, that these are gross elasticities including an income effect, and that many of these elasticities are not significantly different than zero in statistical terms. To test the hypothesis that elasticity values were unaffected by structural change, the standard errors of elasticity differences between the 2 regimes were computed. None of these individual elasticities turned out to be statistically different between the two regimes, and the Wald test reported in Table 4 (and distributed as χ^2 with 9 degrees of freedom) shows

Table 4 - Marshallian Elasticities at the Sample Mean

Elasticity of	Price of Beef	Price of Pork	Price of Chicken	Price of Fish	Expenditure
			Before Structural Change		
Beef	-0.983 (0.068)	-0.004 (0.038)	-0.124 (0.016)	-0.109 (0.023)	1.220 (0.083)
Pork	0.087 (0.119)	-1.015 (0.075)	-0.047 (0.029)	-0.066 (0.040)	1.041 (0.148)
Chicken	-0.161 (0.147)	0.086 (0.092)	-0.090 (0.061)	-0.073 (0.073)	0.238 (0.200)
Fish	-0.182 (0.143)	-0.021 (0.094)	-0.092 (0.054)	-0.138 (0.140)	0.432 (0.247)
			After Structural Change		
Beef	-1.050 (0.064)	-0.078 (0.041)	-0.129 (0.018)	-0.138 (0.021)	1.394 (0.093)
Pork	0.134 (0.101)	-0.839 (0.072)	-0.093 (0.032)	-0.054 (0.033)	0.853 (0.153)
Chicken	-0.017 (0.114)	-0.068 (0.084)	-0.104 (0.072)	-0.022 (0.059)	0.211 (0.172)
Fish	-0.098 (0.110)	0.012 (0.077)	-0.032 (0.055)	-0.196 (0.084)	0.314 (0.165)

Note: Asymtotic standard errors are reported in parentheses;
Wald test = 11.88 .

that they are not significantly different as a set either.

The above results characterizing the elasticity effect of structural change are further illustrated in Table 5, which reports the Allen elasticities of substitution, again evaluated before and after the structural change. The elasticities of substitution that reflect the effects of structural change are defined as:

$$(9.1) \qquad \sigma_{ii} = (\beta_{ii}+\delta_{ii})/(w_i^a)^2 - 1/w_i^a + 1$$

$$(9.2) \qquad \sigma_{ij} = (\beta_{ij}+\delta_{ij})/(w_i^a w_j^a) + 1$$

and, as before, elasticities before structural change can be obtained by setting the δ parameters in (9) to zero, and substituting w_i^b for w_i^a. Table 5 shows that there are not any complementarity relations that are statistically significant. The strongest substitutability relationships appear to exist between beef and pork in both regimes, and between pork and chicken before structural change. After structural change, a subtitutability relationship is

Table 5 - Elasticities of Substitution at the Sample Mean

	Beef	Pork	Chicken	Fish
		Before Structural Change		
Beef	-0.620 (0.083)	1.205 (0.136)	-0.063 (0.165)	0.092 (0.197)
Pork		-2.682 (0.276)	0.552 (0.301)	0.356 (0.338)
Chicken			-0.690 (0.694)	-0.518 (0.635)
Fish				-0.995 (1.257)
		After Structural Change		
Beef	-0.684 (0.080)	1.118 (0.119)	0.178 (0.148)	0.121 (0.158)
Pork		-2.137 (0.221)	-0.032 (0.261)	0.356 (0.244)
Chicken			-0.774 (0.661)	0.010 (0.494)
Fish				-1.492 (0.723)

Note: Asymtotic standard errors are reported in parentheses;
Wald test = 6.05 .

more apparent between beef and chicken. None of these elasticities, however, are statistically different between the two regimes. The Wald test in Table 5 (distributed as χ^2 with 6 degrees of freedom) shows that these elasticities, as a group, are not statistically different between the two regimes.

Conclusions

The evidence presented in this paper supports the idea that the observed meat consumption patterns of the last twenty years cannot be fully explained by the dynamics of prices and income, as the hypothesis of constancy of the parameters of a reasonably specified AIDS model for four meats was rejected against a more general time varying parameter model. This finding was further qualified by the estimated distorting effects of structural change. This showed that the time varying component of the model is biased against beef and in favour of chicken and fish. While the analysis of this paper cannot offer much insight as to the roots of this bias, this movement towards an increased importance of white meats lends further support to the idea that dietary concerns are partly responsible for the perceived changes in meat consumption patterns. The implications of this are particularly relevant to the beef industry, calling possibly for a quality adjustment in production and increased efforts in promotion and marketing.

Evidence of a changing structure of meat demand also has implications for the use of econometric models for forecasting purposes, and this is irrespective of whether the structural change is taken as evidence of changing tastes and preferences or of some other model misspecification. For forecasting purposes, what is called for is a model that can adequately deal with the changing structure. For this reason this paper has attempted to identify the timing and shape of structural change. As the results indicate a fairly rapid transition to a new demand regimes in the mid-1970's, a possible solution would be the use of time series data pertaining to the post 1976 period. Some caution is however warranted, as it was pointed out that a variety of time paths for the structural change cannot be ruled out by the model, including paths involving parameter change up to the end of the period.

Notes

1 For a comprehensive review of meat demand studies, including studies on structural change, see Smallwood, Haidacher and Blaylock (1986).

2 Since the adding up conditions (2.2) will be satisfied by definition by the estimating model, it is clear that the symmetry constraints (2.3) will always imply the remaining homogeneity restrictions (2.1).

3 Note that the properties of homogeneity and symmetry require a set of restrictions on the γ and δ parameters similar to those in (2).

4 It is apparent that the approach of this paper has much in common with the gradual switching regression model of Tsurumi (1982), also applied to structural change in meat demand by Dahlgran (1987). The main difference lies in the choice of the transition function for the path of structural change, and in the fact that we allow structural change to affect all parameters simultaneously (which is consistent with the lack of any strong prior on the characteristics of the change).

REFERENCES

Barten, A. P. "Maximum Likelihood Estimation of a Complete System of Demand Equations." European Economic Review, 1(1969):7-73.

Binswanger, H. "The Measurement of Technical Change Biases with Many Factors of Production." American Economic Review, 64(1974):964-976.

Box, G.E.P. and Jenkins, G.M. Time Series Analysis, Prediction, and Control. San Francisco:Holden-Day, 1970.

Braschler, C. "The Changing Demand Structure for Pork and Beef in the 1970's: Implications for the 1980s." Southern Journal of Agricultural Economics, 15(1983):105-110.

Chalfant, J.A. and Alston, J.M. "Accounting for Changes in Tastes." Journal of Political Economy, 96(1988):391-410.

Chalfant,J.A. "A Globally Flexible, Almost Ideal Demand System." Journal of Business and Economic Statistics, 5(1987):233-242.

Chavas, J.P. "Structural Change in the Demand for Meat," American Journal of Agricultural Economics, 65(1983):148-153.

Chavas, Jean-Paul. "On the Structure of Demand for Meat." Paper presented at a Symposium for Meat, Charleston, South Carolina, 1986.

Chow, G.C. "Tests of Equality between Subsets of Coefficients in Two Linear Regressions." Econometrica, 28(1960):591-605.

Christensen, L. R., and Manser, M. E. "Estimating U.S. Consumer Preferences for Meat with a Flexible Utility Function." Journal of Econometrics, 5(1977):37-53.

Dahlgran, R.A. "Complete Flexibility Systems and the Stationarity of U.S. Meat Demand." Western Journal of Agricultural Economics, 12(1987):152-163.

Deaton, A. "Demand Analysis." In Z. Griliches and M.D. Intriligator, eds., Handbook of Econometrics, Volume III. New York: Elsevier Science Publishers, 1986.

Deaton, A., and Muellbauer, J. "An Almost Ideal Demand System." American Economic Review, 70 (1980):312-326.

Engle, R. F. "Wald, Likelihood Ratio, and Lagrange Multiplier Tests." In Z. Griliches and M. D. Intriligator, eds., Handbook of Econometrics, Volume II. New York: Elsevier Science Publishers, 1984.

Farley, J. V., and Hinich, M. "Testing for a Shifting Slope Coefficient in a Linear Model." Journal of the American Statistical Association, 65(1970):1320-1329.

Hendry, D. F., and Mizon, G. E. "Serial Correlation as a Convenient Simplification, not a Nuisance: A Comment on a Study of the Demand for Money by the Bank of England." Economic Journal, 88(1978):549-563.

Johnson, S.R., Hassan, Z.A., and Green, R.D. Demand System Estimation. Ames: Iowa State University Press, 1984.

Moschini, G., and Meilke, K. D. "Parameter Stability and the U.S. Demand for Beef." Western Journal of Agricultural Economics, 9(1984): 271-282.

Nyankori, J.C.O. and G.H. Miller. "Some Evidence and Implications of Structural Change in Retail Demand for Meats," Southern Journal of Agricultural Economics, 14(1982):65-70.

Ohtani, K., and Katayama, S. "A Gradual Switching Regression Model with Autocorrelated Errors." Economics Letters, 21(1986):169-172.

Smallwood, D.M., Haidacher, R.C. and Blaylock, J.R. "A Review of the Research Literature on Meat Demand." Paper presented at a Symposium for Meat, Charleston, South Carolina, 1986.

Smith, T.K. "Changing Tastes: By end of this year, Poultry will Surpass Beef in the U.S. Diet." The Wall Street Journal, Sept. 17, 1987.

Thurman, W. N. "The Poultry Market: Demand Stability and Industry Structure." American Journal of Agricultural Economics, 69(1987):30-37.

Tsurumi, H. "A Bayesian and Maximum Likelihood Analysis of a Gradual Switching Regression in a Simultaneous Equation Framework." Journal of Econometrics, 19(1982):165-182.

Wohlgenant, M.K. "Estimating Cross Elasticities of Demand for Beef." Western Journal of Agricultural Economics, 10(1985):322-329.