

Measuring an Almost Ideal Demand System with Generalized Flexible Least Squares

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Abstract

Structural change in meat consumption has been the focus of many researchers during the last two decades. In this paper we develop a dynamic linear Almost Ideal Demand System (AIDS) model from a cost function that allows for time varying parameters. This model is consistent with inertia in the parameters of the cost and indirect utility functions. It allows for persistent preferences which may arise from cultural biases, lifestyles, peer pressure, etc. An empirical application is conducted with US meat consumption and price data using a generalized system of flexible least squares, Generalized Flexible Least Squares (GFLS). GFLS allows parameters to evolve slowly over time through incorporating of penalties in fluctuations. Estimated quarterly elasticities were subjected to additional analysis to determine how highly they were related to the Brown and Schrader Cholesterol Index and relative prices. The combined results support that the movements of elasticities over time are related to both.

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Per capita beef consumption in the United States today is nearly the same as it was in the mid-1960's. The interim period, however, saw a 25 percent increase between 1967 and 1976 followed by a 23 percent decrease from 1976 to 1988. Attempts to explain these wide swings in meat consumption patterns over the past 25 years have used a variety of techniques and functional forms. The debate that has evolved is much like the popular lite beer commercial where one group chants "tastes great" and the other responds "less filling". The two cheers heard in the meat demand arena are "relative prices" and "changed preferences". Of course the purpose of the beer commercial is to persuade consumers that the product is less filling and tastes great. Likewise, this paper attempts to present an analysis that is capable of encompassing an important role for both relative prices and changing consumer tastes in the context of meat demand.

With the exception of Chavas, the previous structural change models have not accounted for the persistence of the parameters of preferences over time. That is, any structural change due to changes in preferences should be expected to occur slowly or smoothly as suggested by Moschini¹. Slowly evolving preferences are consistent with the effects of cultural biases, lifestyles, peer pressure, and habit formation, all of which prohibit rapid adjustments in utility function parameters. The switching regression models estimated by Braschler and Moschini and Meilke (1989) generated fairly discrete changes in parameter values. Both papers identified significant structural change in the mid-1970's. The results do not provide a means to determine the relative importance of preference changes and relative prices. The Moschini and Meilke (1989) paper estimated a statistically significant linear trend in the parameters of an almost ideal demand system for the four quarters between 1975-IV to 1976-III. For the beef industry, and by association other livestock, this was a fairly aberrant year. It follows closely on the heels of

price controls that had a profound impact on farm level cattle prices and the supply of cattle. Many farmers held cattle back from slaughter hoping to capitalize on anticipated higher prices when the price controls were lifted. The effect was an over-supply of cattle, a drastic drop in cattle prices, and the beginning of a sustained liquidation of the aggregate cattle herd. One aspect of the available aggregate data is that beef is a perishable commodity and that prices often adjust to clear whatever quantity is produced. Consequently, it is possible that the mid-70's represent a phenomena best handled by a dummy variable rather than a permanent change in structure.

Using different functional forms Moschini and Meilke (1984) and Wohlgenant failed to find significant structural change in meat demand. Wohlgenant concluded that previous evidence of structural change was confounded with the specification of functional form. Chalfant and Alston used a nonparametric test to determine that meat demand data were consistent with unchanged tastes over time. In another paper, Alston and Chalfant show that Canadian meat demand data pass the nonparametric test, however, the same data revealed significant structural change within the context of two parametric forms.

Time-Varying Consumer Preferences

As mentioned above many authors have generated plausible models with either dynamic coefficients or lagged variables (see Blanciforti and Green, and Eales and Unnevehr for the latter). In order to introduce time varying coefficients into a demand system some theoretical, as well as heuristic arguments are valuable. To begin with, consider a world with intertemporally additive preferences represented by the PIGLOG cost function.

$$\ln c(u, p) = (1 - u) \ln a(p) + u \ln b(p) \quad (1)$$

¹Moschini used a time trend and the Brown and Schrader cholesterol index to mimic smooth preference changes.

where $a(p)$ and $b(p)$ are both homogeneous of degree one in prices, p , and u is utility ranging from zero to one. This is familiar as the cost function which Deaton and Muellbauer used to develop the almost ideal demand system². Given Deaton and Muellbauer's parameterizations and the cost function in (1) it is possible to solve for the indirect utility function

$$u = \frac{\ln(x/a(p))}{\beta_0 \prod_k p_k^{\beta_k}} \quad (2)$$

where x is total expenditure and is equal to $c(u,p)$ for the utility maximizer.

Typically, structural change is considered to involve changes in parameter values within a single specification. In the context of the almost ideal demand system this implies that the parameters of $\ln a(p)$ and $\ln b(p)$ change over time. Consequently, alternative specifications of $\ln a(p)$ and $\ln b(p)$ which are capable of allowing a wide scope of parameter mobility are needed. Chavas has indirectly implied a form for these new specifications in his earlier Kalman filter paper. Consider the following

$$\ln a(p) = \alpha_{0t} + \sum_k \alpha_{kt} \ln p_{kt} + \frac{1}{2} \sum_k \sum_j \gamma_{jtk} \ln p_{jt} \ln p_{kt} \quad (3)$$

with

$$\phi_{t+1} = F(t)\phi_t + e_t$$

Where ϕ is a vector of parameters $\{\alpha_0, \dots, \alpha_k, \dots, \gamma_{11}, \dots, \gamma_{kk}\}$. The $F(t)$ matrix is the dynamic equation matrix, and e_t is a random forcing term. Further Chavas (1983) specifies

$$\ln b(p) = \ln a(p) + \beta_0 \prod_k p_{kt}^{\beta_{kt}} \quad (4)$$

² The choice of $\ln a(p)$ and $\ln b(p)$ determine the characteristic of the demand equations, as well as, the flexibility of the cost function. Deaton and Muellbauer chose the following forms

$$\ln a(p) = \alpha_0 + \sum_k \alpha_k \ln p_k + \frac{1}{2} \sum_k \sum_k \gamma_{jk} \ln p_j \ln p_k$$

with

$$\beta_{k,t+1} = G(t)\beta_{k,t} + u_{k,t} \quad \forall k$$

Together these specifications allow the own and cross price elasticities to vary over time with preferences, as well as, price and quantity levels. The result is a linkage of consumption patterns over time. Thus, the intertemporal separability assumption is somewhat weakened since estimates of ϕ_{t+1} and β_{t+1} will, in general, be calculated as functions of past quantities and prices.

To further develop the model we use equations (1) - (4) and Shepard's Lemma to derive a system of share equations which resemble the usual static almost ideal demand system, but have the innovation that the parameters move according to the time paths described for them above. The i^{th} share equation is given by

$$w_{it} = \alpha_{it} + \sum_j \gamma_{ijt} \ln P_{jt} + \beta_{it} \ln \left(\frac{x_t}{P_t} \right) \quad (5)$$

where P_t is the Stone's price index commonly used to linearize the almost ideal demand system, and x_t is total expenditures on meat. The parameters of this demand system evolve from their predecessors over time taking into account the inertia in human behavior, and because of the random forcing terms, can represent any form of structural change if it occurs. Undoubtedly, constant changes in preferences are occurring, but because the observed data is highly aggregated it is difficult to model, with standard techniques, anything except very large shifts or an accumulation of small changes in the same direction. This demand system assumes that smaller adjustments can be accurately modeled with existing data. The analysis with switching regimes where either a single shift or constant linear trends were found suggests simplistic regimes rather than a perpetually evolving system of preferences. With the above analysis, the

$$\ln(b(P)) = \ln a(P) + \beta_0 \prod_k P_k^{\beta_k}$$

parameters could move according to any of a variety of dynamic patterns including, logistical, random, or linear trend.

Few supporters of a preference change in meat consumption would disagree that such changes occurred over a long period of time between the 1970's and present. In fact, the adjustments may not be complete. As prices, nutritional information, and product characteristics evolve in a dynamic economy decisions made in the past continue to influence the supply and demand for meat and other products.

In previous meat demand studies significant parameter changes have been discovered. However, Wohlgenant demonstrated that such statistical results can be dependent on the specification of the demand system used. Therefore, it is insufficient to simply suggest a time-varying parameter model and estimate it. Time-varying elasticities can be calculated from the above almost ideal demand system and these elasticities could vary for two main reasons. First, due to relative prices, and second, due to preference changes. If the structural changes implied by the parameter estimates are the result of an incorrect specification, then they should not vary systematically with any measure of consumer tastes. That is, the elasticities should not be significantly affected by such a measure, but should vary significantly only with relative price changes. A straightforward test of the importance of these factors can be implemented by regressing the elasticities against relative prices and a measure of consumer tastes.

Data

The data used in this analysis are identical to those used by Moschini and Meilke (1989). They consist of quarterly per capita disappearances and retail prices of beef, chicken, fish and seafood, and pork. Quantities and prices for chicken are those published by USDA in *Poultry and Egg Situation* and *Livestock and Poultry Situation*. The quantity reflects per capita

disappearance of total chicken (young and mature), and the price is that of fryers. Choice beef and pork quantities and prices came from USDA's *Livestock and Meat Situation*. Fish and seafood consumption data are derived from USDC data on personal expenditures and the CPI for fish and seafood. The sample begins in 1966 and extends through 1987 spanning the mid-1970's, the period of structural change discovered by previous research. The Brown and Schrader cholesterol index was chosen to represent potential change in consumer tastes. This index is the sum of the number of medical journal articles supporting a link between cholesterol intake and heart disease minus those that do not support such a link. The articles were identified using the Medline database. Brown and Schrader suggest using the two-quarter lag of this sum in order to capture the lag time between publication of the article and the impact (via physicians) on consumers. While this represents a fairly ad hoc distributed lag effect it was deemed satisfactory for this study since the purpose is to demonstrate a significant link to changing meat consumption not accurately quantify the marginal impact of another medical journal article.

Empirical Implementation

Adding a disturbance term to the share equations in (5) yields an empirical model of the type which West and Harrison refer to as a Dynamic Linear Model (DLM). In general terms the DLM is written

$$y_t = x_t \beta_t + u_t \quad (6a)$$

$$\beta_{t+1} = F_t \beta_t + e_t \quad (6b)$$

Equation (6a) is commonly called the observation equation and (6b) is referred to as the system evolution equation.

There are a number of related methods that can be used to estimate the unknown parameters, β_t . In fact, the system estimated by Chavas is similar although his specification is

slightly different from the almost ideal demand system posited in the previous section. He used a standard Kalman filter to estimate time updated parameter values and found structural change in beef and chicken consumption in the 1970's. Bayesian techniques could also be used if workable priors can be identified.

The approach chosen in this paper is in a sense a hybrid of the Kalman filter and Bayesian estimation. The Generalized Flexible Least Squares (GFLS) approach is a multiple equation extension of the Flexible Least Squares (FLS) technique developed by Kalaba and Tesfatsion. In fact, Dorfman and Foster used the single equation FLS model to estimate technical change in a production economics setting. GFLS represents a special case of the Kalman filter in which updates in parameter estimates are penalized for deviations from previous values. This slows and smooths parameter movement over time and is consistent with the notion that preferences tend to be sticky³.

The GFLS algorithm seeks to minimize a loss function which includes both the usual sum of squared errors from the observation equation (SSE), and the sum of squared dynamic errors from the evolution equation (SSD). For the GFLS case the loss function is

$$C(\beta; \mu, T) = \mu \sum_{t=1}^{T-1} [\beta_{t+1} - F_t \beta_t]' D_t [\beta_{t+1} - F_t \beta_t] + \sum_{t=1}^T [y_t - x_t \beta_t]' M_t [y_t - x_t \beta_t] \quad (7)$$

GFLS solves for the trajectory of β that minimizes the loss function over time. The matrix M_t is analogous to the variance-covariance matrix in standard generalized least squares regression. Under the assumption of contemporaneously correlated equation disturbances the ij^{th} element of M is $u_i' u_j / (n - k)$. The parameter μ defines the relative importance of the two terms in the objective functions. When values of μ approach infinity the GFLS estimates converge to

constant parameter seemingly unrelated regressions estimates, because μ represents the penalty for parameter movements over time. Setting μ equal to zero leads to a random coefficients model analogous to that developed by Singh and Ullah, because there are no penalties for parameter movement⁴. As a consequence, there is a trade-off between SSE and SSD which can be depicted graphically as in Dorfman and Foster's efficiency frontier when the value of μ is varied from 0 to ∞ . In the classical econometric context μ is the inverse of the prior precision for a stochastic restriction. Thus, in past applications the specification of μ has been arbitrary.

In order to find a less ad hoc value of μ one could begin with the prior belief of a strict classical econometrician that both sums of squares will be zero. That is, the parameters are not time-varying, and the model is the correct specification with zero measurement error and no omitted variables. While few econometricians actually hold this naive view of the world, it remains true that this is the prior belief underlying the majority of applied econometric analyses. The point on the efficiency frontier closest to the origin yields the estimate of μ most consistent with this naive prior belief. Thus, we propose to choose μ by minimizing the vector norm, in the efficiency frontier space, that is defined as

$$J = \sqrt{\frac{SSE_{fls}^2}{SSE_{ols}^2} + \frac{SSD_{fls}^2}{SSE_{RC}^2}} \quad (8)$$

where SSE_{ols}^2 is the sum of squared errors when the model is estimated with constant parameters (i.e. SUR), and SSE_{rc}^2 is the sum of squared errors when the model is estimated as a seemingly unrelated random coefficients model, and the *fls* subscripts imply the counterparts when the

³The reference to slow and smooth parameter dynamics should not mislead one to believe that GFLS is incapable of allowing for rapid structural shifts. Kalaba and Tesfatsion and Tesfatsion and Veitch have demonstrated the ability of FLS estimates to track a wide variety of "smooth" and "nonsmooth" (rapid) parameter transitions.

⁴ Specifying the matrix D as other than an identity matrix will result in separate penalties for movement of the elements in β_t over time.

model is estimated by generalized flexible least squares. The scaling is necessary to prevent the relative size of the dynamic and measurement errors from skewing the efficiency frontier. The value of μ chosen by this method represents an interpretable specification in the context of the constant parameter model in two ways. First, the estimates can be compared to the constant parameter estimates. Second, it is the model that the data suggests is most consistent with both constant parameters and a zero SSE.

We maintained the assumptions of weak separability between meat and other commodities and intertemporal separability to simplify the estimation and remain consistent with previous meat demand studies. Furthermore, the dynamic equation matrix was assumed to be an identity. The weighting matrix D in equation (7) was held constant over time. The off-diagonal elements of this matrix were set to zero and the diagonal elements associated with independent variables other than quarterly dummies were set to unity. The diagonal elements associated with the quarterly dummy variables were set to ten million in order to prevent any movement in their parameter estimates. Quarterly dummy variables with time-varying coefficients could potentially pick up all of the variation in the shares and lead to unsatisfactory estimates of the remaining coefficients.

To implement this estimation a search routine was used to find the value of μ which minimized the norm defined in equation (8). The routine parametrically varied μ from zero to 19,035⁵ to minimize J . The value of μ that minimized J was found to be 2.14. This appears to be closer to the random coefficients model than the SUR model and is close in magnitude to the value used by Dorfman and Foster. Figures 1 and 2 show the movement of the own price coefficients estimated by the GFLS/AIDS approach. Notice that both the beef and chicken

⁵ 19,035 was the largest value of μ that produced non-zero estimates for the coefficients.

coefficients follow upward trends, with the own price coefficient for chicken increases more than four times from 1974-I to 1975-I. The own price coefficient for pork falls until 1980-81 and then increases over the remaining periods in the sample to higher than initial levels. The own price coefficient for fish trends down over the entire sample with some large changes between 1974 and 1976. Table 1 presents the mean, minimum, maximum, and standard deviations of all parameter estimates. The model fit was very satisfactory. For all equations in the system, the squared correlation between actual and predicted expenditure shares exceeded 0.98.

Another interesting artifact of the chosen estimation method is the opportunity to evaluate the bias in structural change over time. Antle and Capablo present a good discussion of the multifactor dual (cost function) measure of bias first derived by Binswanger (1974, 1978). Furthermore, Antle and Capalbo show that Binswanger's bias measure contains a confounding scale effect when technology (preference) is not homothetic. They suggest the correct scale adjustment that leaves a pure measure of bias. An analogous approach works here since the GFLS/AIDS model is derived from the expenditure equation (1). This measure is defined as⁶

$$B_{it} = \frac{\partial \ln w_{it}}{\partial t} - (\eta_{it} - 1) \frac{\partial \ln c_t}{\partial t} \quad (9)$$

where η_{it} is the expenditure elasticity of demand and B_{it} is the scale adjusted bias for the i^{th} good in quarter t . The biases were calculated at the estimated shares for each quarter, but because the parameter u in equation (1) is unobservable it was necessary to use actual expenditures in forming bias estimates.

⁶The equation presented by Antle and Capalbo is slightly different. They write:

$$B_i = \frac{\partial \ln w_i}{\partial t} - \frac{\partial \ln w_i}{\partial Q} \left(\frac{\partial \ln C}{\partial Q} \right)^{-1} \frac{\partial \ln C}{\partial t}$$

However, replacing Q with u, differentiating equations (1) and (5), and substituting into the above formula yields equation (9).

The estimated structural change biases are shown in Figures 3 and 4. The appropriate interpretation is of importance. Positive values imply that the structural change favors that factor. A glance at the figures shows that the biases fluctuate greatly over the sample. However fish biases have tended to be positive while beef, pork, and chicken have tended to be negative throughout the sample. The biases for fish and pork have been relatively small over the sample compared to beef and fish biases. The mean biases before 1976 are basically unchanged versus those for the post 1976 period. The means for the entire sample are -0.0004, -0.0021, -0.0005, 0.0001 for beef, pork, chicken, and fish respectively. While these are small, the graph shows that for individual sample periods some of the biases are large⁷. Moschini and Meilke (1989) found somewhat different aggregate biases. In the cases of beef only the mean GFLS/AIDS bias has a compatible sign with the Moschini and Meilke (1989) estimates. Only the pork bias estimates are similar in magnitude. These differences may arise from the scale effect adjustments in the GFLS/AIDS bias estimates or from differences in estimation techniques. The GFLS/AIDS results suggest that, with the exception of fish, the periods of negative bias have outweighed the positive.

The estimated elasticities derived from the GFLS/AIDS model are of also of interest in evaluating how meat demand has evolved over time. The uncompensated elasticities were calculated for each quarter (at the mean of the predicted shares) with the following formula suggested by Green and Alston,

$$E = [BC + I]^{-1}[A + I] - I \quad (10)$$

where the typical elements are: $a_{ij} = -\delta_{ij} \left(\frac{\gamma_{ij}}{w_i} \right) - \beta_i \left(\frac{w_j}{w_i} \right)$ in A ; $b_i = \left(\frac{\beta_i}{w_i} \right)$ in B ; $c_j = w_j \ln p_j$ in C ; ϵ_{ij} in E ; and δ_{ij} is one if $i=j$ and zero otherwise. This formula is slightly

⁷The means of the absolute values of the biases were 0.0086, 0.0156, 0.060, and 0.0030.

different from that used by Chalfant and others but it has been shown, in Monte Carlo experiments, to outperform the more traditional alternatives (see Alston, Foster, and Green).

To conserve space, only the own price elasticities are presented in Figures 5 and 6. Figure 5 displays the time path of the own price elasticities of beef and chicken demand⁸. Notice that the own price elasticity of beef and chicken trends upward or becomes less responsive to price changes, with chicken experiencing a large decrease in responsiveness around 1974-1975. To the contrary, the pork demand own price elasticity increases in absolute value or becomes more responsive to price changes, while the fish own price elasticity increases in the later portion of the sample. In the case of beef, the absolute change over the sample period is not that large but is enough to have meaningful economic consequences. These elasticity estimates differ from those of Moschini and Meilke (1989). They found that beef demand was more responsive and pork demand was less responsive after the mid-1970's structural change and that chicken demand did not change much. However, the GFLS/AIDS estimates for fish own price elasticities are in agreement with those of Moschini and Meilke (1989), that is both suggest that the own price elasticity for fish has become more responsive. Consistency between the two sets of results is also mixed for the uncompensated cross-price elasticities. The GFLS/AIDS elasticity estimates with the same sign and direction of change as those estimated by Moschini and Meilke (1989) were: ϵ_{bp} , ϵ_{bc} , ϵ_{bf} , ϵ_{cb} , ϵ_{fb} , ϵ_{fp} , and ϵ_{fc} . The GFLS/AIDS estimates for ϵ_{bp} and ϵ_{fp} were similar in sign but less pronounced than Moschini and Meilke's (1989) results.

Determinants of Structural Change

In the previous sections, changes in the own and cross-price demand elasticities gave evidence of changing consumption patterns. In this section, an attempt is made to determine the

significant causes of these changes. Two schools of thought have been suggested: changes in preferences and changes in relative prices. These hypotheses suggest two competing specifications for explaining intertemporal variations in the price elasticities of demand for meat. The elasticities used here are slightly different from those discussed in the previous section. In order to test these hypotheses the following non-nested models were specified.

Relative Prices Model:

$$\varepsilon_{ij} = \alpha_0 + \alpha_b \frac{P_b}{P_c} + \alpha_p \frac{P_p}{P_c} + \alpha_f \frac{P_f}{P_c} + \alpha_D D + u_{ij} \quad (11)$$

Cholesterol Index Model

$$\varepsilon_{ij} = \beta_0 + \beta_{CH} CH + \beta_D D + v_{ij} \quad (12)$$

where the α_i and β_i are parameters to be estimated, CH is the Brown and Schrader cholesterol index divided by its mean, D is a matrix of quarterly dummy variables for the first three quarters, and the u_{ij} and v_{ij} are contemporaneously correlated disturbance terms. Note that since all equations in each set (Relative Prices Model and Cholesterol Index Model) contains the same regressors it is sufficient to estimate each of the sixteen equations of the form in (11) and (12) by ordinary least squares⁹.

Significant autocorrelation was detected in each of the equations in (11) and (12). This along with the quarterly dummy variables presents a difficulty with respect to a straightforward application of the J-test. The original development by Davidson and McKinnon did not address the possibility of autocorrelation. Anatonovitz and Green point out that the correction for autocorrelation should be made during the estimation of the competing models, but that this

⁸All of the estimated own price elasticities (compensated and uncompensated) were negative for all quarters. Expenditure elasticities were also calculated. In addition, all of the estimated expenditure elasticities were positive for all quarters.

⁹ SUR becomes OLS when $X_i = X$ for all equations.

information should not be included when calculating the predicted values used in the second step of the J-Test procedure so as not to bias the test in favor of that model. The same can be said of the quarterly dummy variables. Consequently, the equations in (11) and (12) were estimated by the Prais-Winsten method to preserve the initial observation. The predicted values were calculated by the general form:

$$\hat{y}_t = \hat{\alpha} + \hat{\beta}x_t \quad (13)$$

rather than

$$\hat{y}_t = \hat{\alpha} + \hat{\beta}x_t - \rho x_{t-1} + \rho y_{t-1} \quad (14)$$

The results of the J-Test for each of the sixteen uncompensated elasticities and the two models are presented in Table 2. Notice that the cholesterol index model is rejected in 13 out of 16 cases. However, the model based on relative prices cannot be rejected in ten of the cases. The only pattern for those that are not rejected seems to be from the elasticities taken with respect to beef and pork, ϵ_{bb} , ϵ_{bp} , ϵ_{bc} , ϵ_{bf} , ϵ_{pb} , ϵ_{pp} , ϵ_{pc} , ϵ_{pf} , ϵ_{cc} , and ϵ_{fc} .

Conclusions

In this paper a demand system with time varying coefficients (GFLS/AIDS) was developed and estimated using quarterly U.S. meat consumption data. Changes in structure were embodied in the movement of parameter estimates over time. Even with an objective function which penalized parameter movement fairly strong, the parameters achieved economically significant changes over the sample period. Estimates of scale adjusted biases in structural change exposed a fluctuating situation which has, on average, been biased against beef and in favor of fish.

Of greater importance, however, was the discovery of non-random movements in the own price elasticities of demand for beef, pork, and fish. In the case of pork, demand has become

more own price responsive while beef and chicken demand has become less responsive. These trends have significant implications for meat industries. Less elastic beef demand suggests that future cattle price cycles will be larger relative to their accompanying quantity cycles, assuming no changes in the supply schedule. In fact, some of this may already be occurring and could help explain the slow cattle herd build up of the early 90's in the face of high prices. The beef industry would appear to be in a good position to take advantage of an increasingly inelastic demand for their product.

In order to determine the source(s) of changing consumption patterns over time, sets of regressions were estimated using the uncompensated elasticities as dependent variables. The independent variables were relative prices, the Brown and Schrader cholesterol index, and quarterly dummy variables. The results strongly support a significant role for relative prices and changing tastes and preferences. There seems to be less support for the relative price hypothesis for chicken and fish demand elasticities and more for the beef and pork cases (see Table 2). In other words, this lite beer, "tastes great and it's less filling!"

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Table 1. Summary of the GFLS Estimated Coefficients and AIDS/GFLS Estimated Own Price Elasticities

Coefficient	Average	Minimum	Maximum
α_b	0.0273	0.0239	0.0290
α_p	0.0140	0.0117	0.0173
α_c	0.0069	0.0047	0.0107
γ_{bb}	0.0163	0.0070	0.0267
γ_{bp}	0.0134	0.0031	0.0198
γ_{bc}	0.0159	0.0040	0.0218
γ_{bf}	-0.0457	-0.0533	-0.0273
γ_{pp}	0.0066	-0.0007	0.0184
γ_{pc}	0.0033	-0.0108	0.0131
γ_{pf}	-0.0234	-0.0413	-0.0082
γ_{pp}	0.0187	0.0030	0.0342
γ_{pf}	-0.0379	-0.0449	-0.0321
γ_{ff}	0.1070	0.0769	0.1230
β_b	0.1109	0.0973	0.1166
β_p	0.0632	0.0582	0.0731
β_c	0.0224	0.0183	0.0310
β_f	-0.1965	-0.2125	-0.1802
ϵ_{bb}	-0.4497	-0.4904	-0.4249
ϵ_{pp}	-1.3807	-1.4541	-1.3419
ϵ_{cc}	-0.9915	-1.1608	-0.8654
ϵ_{ff}	-0.1407	-0.1511	-0.1320

^a Subscripts b, p, c, and f denote beef, pork, chicken, and fish.

Table 2. J-Test Results for Relative Prices and Cholesterol Index Models of GFLS/AIDS Uncompensated Demand Elasticities

Elasticity	Coefficients from the Relative Prices Model			Results of J-Test at 5% Level	
	α_b	α_p	α_f	Relative Prices Model	Cholesterol Model
ϵ_{bb}^b	0.009* ^c (0.003) ^d	0.008* (0.006)	0.011 (0.002)	<i>can't reject</i>	<i>reject</i>
ϵ_{bp}	-0.003* (0.001)	-0.002 (0.001)	-0.001* (0.001)	<i>can't reject</i>	<i>reject</i>
ϵ_{bc}	-0.002 (0.001)	-0.001 (0.002)	-0.003* (0.001)	<i>can't reject</i>	<i>reject</i>
ϵ_{bf}	-0.006* (0.002)	-0.006 (0.003)	-0.004* (0.001)	<i>can't reject</i>	<i>reject</i>
ϵ_{pb}	-0.022* (0.003)	-0.008 (0.005)	0.009* (0.002)	<i>can't reject</i>	<i>can't reject</i>
ϵ_{pp}	0.026* (0.009)	-0.082 (0.015)	-0.011* (0.004)	<i>can't reject</i>	<i>can't reject</i>
ϵ_{pc}	-0.035* (0.007)	-0.008 (0.011)	0.022* (0.003)	<i>can't reject</i>	<i>reject</i>
ϵ_{pf}	-0.021* (0.003)	-0.004 (0.006)	0.011* (0.002)	<i>can't reject</i>	<i>reject</i>
ϵ_{cb}	-0.016* (0.002)	-0.009* (0.003)	0.007* (0.001)	<i>reject</i>	<i>reject</i>
ϵ_{cp}	-0.009* (0.002)	-0.011* (0.004)	0.010* (0.001)	<i>reject</i>	<i>reject</i>
ϵ_{cc}	0.009 (0.030)	0.181* (0.051)	0.036* (0.015)	<i>can't reject</i>	<i>reject</i>
ϵ_{cf}	-0.027* (0.005)	-0.037* (0.008)	0.037* (0.002)	<i>reject</i>	<i>reject</i>
ϵ_{fb}	-0.005* (0.0001)	-0.003* (0.0001)	0.002* (0.0001)	<i>reject</i>	<i>can't reject</i>
ϵ_{fp}	-0.003* (0.0001)	-0.002* (0.0001)	0.002* (0.0001)	<i>reject</i>	<i>reject</i>
ϵ_{fc}	-0.001* (0.0001)	-0.001 (0.001)	0.001* (0.0003)	<i>can't reject</i>	<i>can't reject</i>
ϵ_{ff}	-0.005* (0.001)	-0.006* (0.002)	0.003* (0.0006)	<i>reject</i>	<i>reject</i>

^b Subscripts b, p, c, and f denote beef, pork, chicken, and fish.

^c * indicates significance at the 5% level.

^d The numbers in parentheses are standard errors.

Figure 1. Estimated Own Price Coefficients from GFLS/AIDS Model for Beef and Chicken

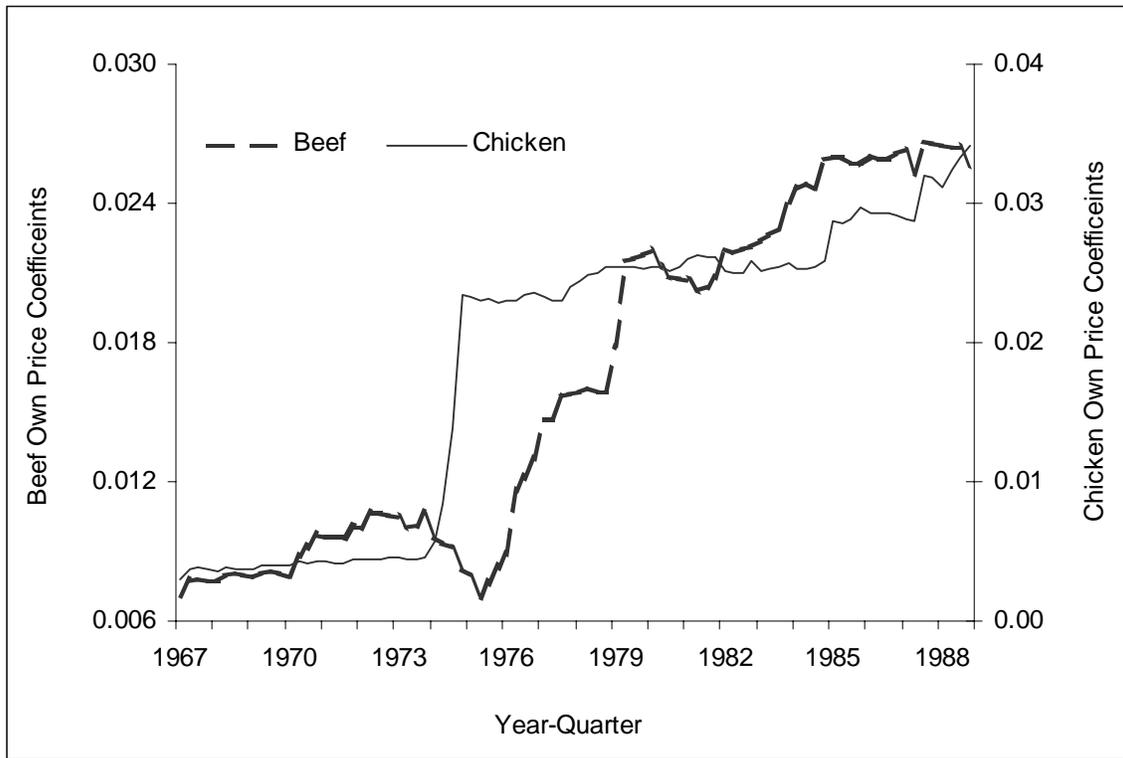


Figure 2. Estimated Own Price Coefficients from GFLS/AIDS Model for Pork and Fish

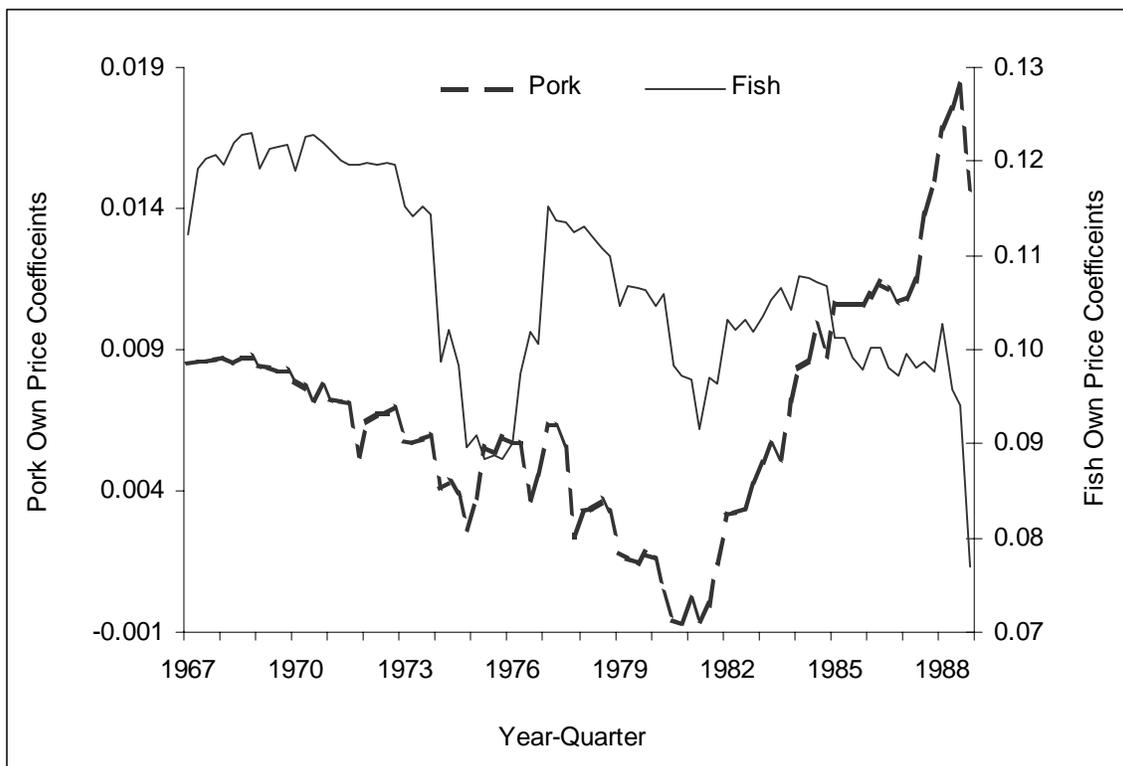


Figure 3. Structural Change Biases in Beef and Chicken Demand

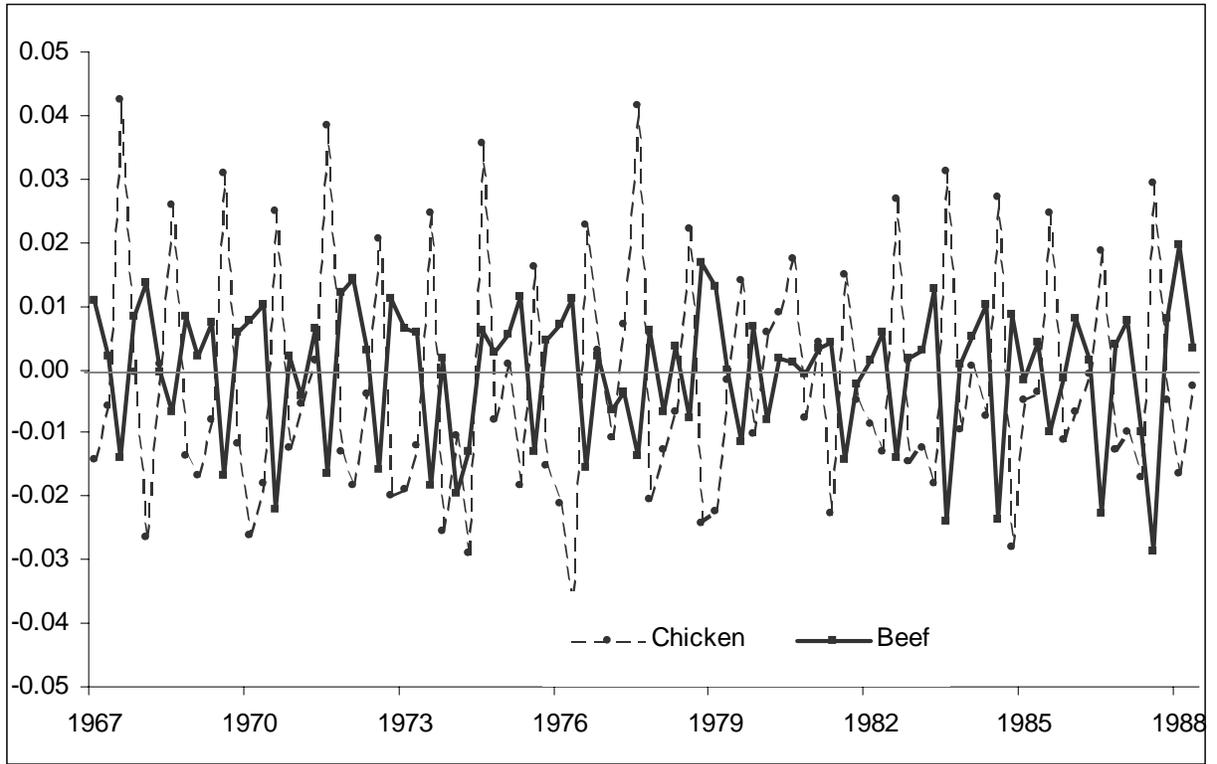


Figure 4. Structural Change Biases in Pork and Fish Demand

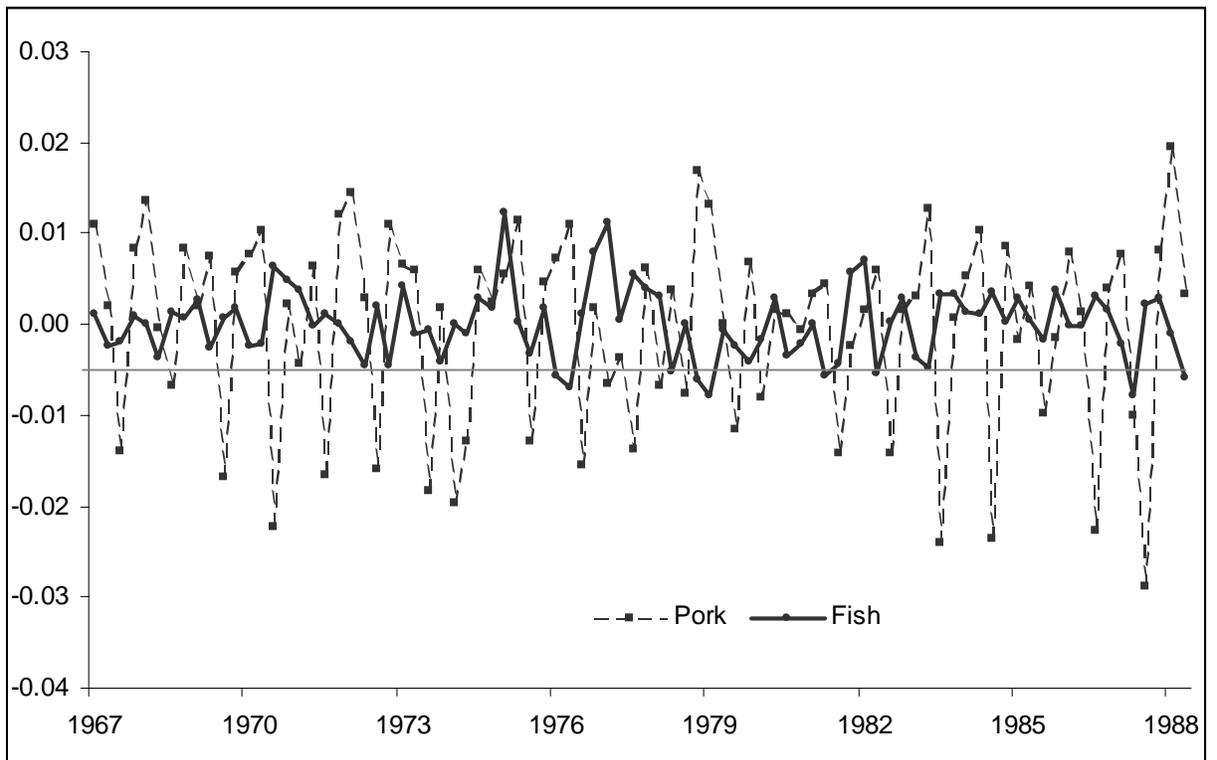


Figure 5. Beef and Chicken Own Price Elasticities

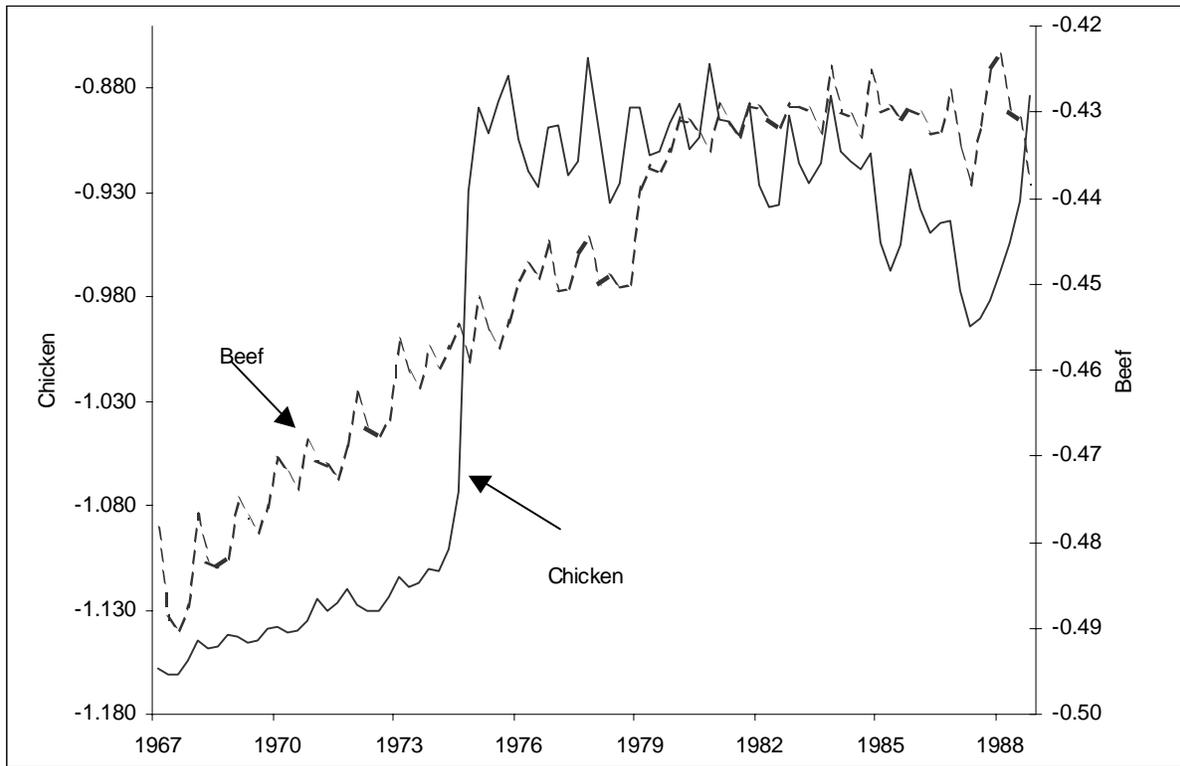


Figure 6. Pork and Fish Own Price Elasticities

