Complete Flexibility Systems and the Stationarity of U.S. Meat Demands

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A Rotterdam demand model is used to detect evidence of structural change in beef, pork, and chicken demands. The demand model is partially inverted prior to estimation to account for meat supply fixity. Estimation uses a likelihood maximization routine applied to 1950 through 1985 annual data. The results suggest severe disruption in the meat markets in the 1970s. A comparison of the 1980s and the 1960s elasticity structures reveals that income and cross-price elasticities are nearly the same but direct price elasticities are lower and are trending toward even more inelasticity. Implications for pricing and risk management are discussed.

Key words: demand systems, flexibility systems, structural change.

Over the past two and one-half decades, U.S. per capita beef and chicken consumption have displayed significant trends, as shown in figure 1. Chicken consumption displays a steady upward trend throughout the period while the general upward trend in beef consumption ended in 1976. Figure 1 also shows the price of beef relative to the price of chicken, which to some extent explains the observed consumption patterns. This explanation assumes that the observed meat consumption patterns are caused by movements in, or fluctuations of, meat prices, consumer incomes, and the prices of substitute goods, all of which interacted with stable meat demand functions. The limited data shown give some credence to this explanation because (*a*) the upward trend in chicken consumption is consistent with the increase in the relative beef price, and (*b*) the drastic increase in the beef-chicken price ratio, starting in 1977, is consistent with depressed beef consumption and the accelerated growth in chicken consumption in the 1980s.

When the news media discuss the changing meat consumption patterns, consumers' blood cholesterol and other nutritional concerns generally receive a great deal of credit for the departures from long-term trends. (For example, see *Business Week,* 26 Aug. 1985, p. 39, and 28 Oct. 1985, p. 40.) These concerns presumably are reflected as a change in the meat demand structure. Although such meat demand changes are not necessary for the observed meat consumption changes, their role as contributing factors may be important.

Empirical meat demand studies have been published by Haidacher et al., Nyankori and Miller, Chavas, Braschler, and Moschini and Meilke. The studies by Nyankori and Miller, Chavas, and Braschler present evidence that structural change occurred in meat demands in the early 1970s, while the studies by Haidacher et al., and Moschini and Meilke find no such evidence. These contradictory results are due to differences in models, data, assumptions, and definitions of structural change.

Because the detection of structural change depends on its definition, structural change must be defined prior to detection. First, as an antidefinition, structural change is not a shift in an empirical demand function when the function excludes the price of either substitute or complementary commodities. Obviously, such a shift could be caused by a change in an excluded price interacting with a stable demand structure. Such errors can be prevented with the use of complete demand systems.

After discussing these issues in greater detail, Haidacher defines structural change in demand as being caused by changes in the rep-

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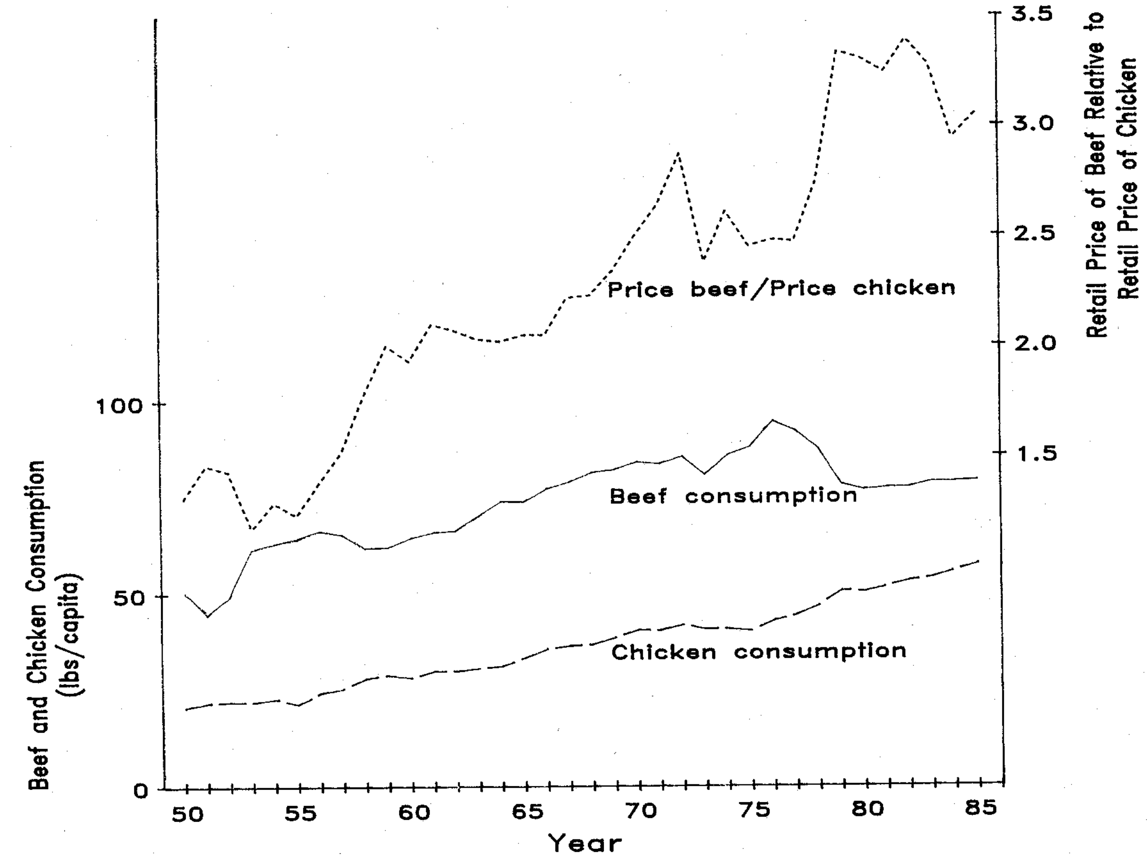

Figure 1. Beef and chicken consumption (pounds per capita) and beef-chicken price ratio. 1950–85

resentative consumer's utility function parameters. Assuming the correctness and completeness of the specified utility function, demand change under this definition can be traced to changes in the representative consumer's utility-function parameters. An alternative definition of structural change, albeit still in the context of complete demand systems, is a change in any of the demand-function parameters, i.e., elasticities or slopes, that describe the decisions of market participants.[1] Assuming the correctness and completeness of the specified demand system under this definition, changes in the demand-function parameters must be caused by changes in consumers' preferences even though the change cannot generally be traced to a specific set of utility function parameters.[2] If the correctness of specification assumption is relaxed, specification error may lead to either erroneous detection or erroneous nondetection of structural change under either procedure. This may occur because the algebraic form of the representative consumer's utility function is not known in applying the first definition and the demand functions are not known in applying the second definition. Hence, invalid assumptions regarding both the constancy of parameters and the

[1] The distinction between these two definitions parallels the distinction in demand theory on the derivation of empirical demand systems. Utility-based demand systems are solutions to a representative consumer's income-constrained maximization of a utility function selected from a family of acceptable alternatives. The linear expenditure system is an example of such a system. An alternative method is to specify directly an empirical demand system which is then restricted to ensure compatibility with the representative-consumer's maximization of a general utility function. The Rotterdam demand system is an example of such a system. For further development of this distinction see Johnson, Hassan, and Green (pp. 62–75).

[2] If the demand system displays the integrability property (Johnson, Hassan, and Green, pp. 35–37), then the underlying utility structure can be derived from the demand system. In this situation, changes in demand structure parameters can be traced to underlying utility parameters.

choice of influencing variables may have been made.

The objective of this paper is to investigate the role attributable to economic variables in the historical changes in meat consumption patterns. This is done by using a complete Rotterdam demand system (Theil 1971, 1975, 1976) to account for the effect of income and price changes on consumption. Should a change in consumers' reactions to economic variables be detected, another objective is to determine whether the change is consistent with a structural demand change, as the news media claims exists, or whether it is due to specification problems. Because the Rotterdam model is derived from per capita demands rather than from the representative consumer's preference structure, the alternative definition of structural change is used.

This study builds on, and extends, previous work on structural change in meat demands by using a complete demand system to search for evidence of structural change. The use of market-level data for estimating the model leads to the assumption of fixed market-level supplies. Prior to estimation, the model is partially inverted to get price-dependent equations for meats while retaining quantity-dependent equations for other foods and nonfood items. Because the model constitutes a complete system, complete elasticity and flexibility matrices (à la Houck 1965, 1966) can be derived.

This modeling approach has merit in that it attempts to reconcile an incompatibility between the data used and demand theory. When demand systems are derived from a postulated utility function (for example George and King; or Green, Hassan, and Johnson), structural change under the first definition is detectable, but it requires assuming that the consumer faces predetermined prices. This study uses per capita consumption data that are derived from market-level disappearance data. In this situation, the assumption of predetermined prices amounts to assuming perfectly elastic market-level supply functions. This assumption is not appropriate for meats because of the relatively long biological gestation and growth processes associated with meat production.

An alternative assumption is to treat market-level, and hence per capita, supplies as fixed. However, simply imposing the assumption of supply fixity on a utility-maximizing representative consumer is not appropriate because supply fixity at the market level does not apply to the individual consumer. Instead, the supply-fixity assumption is imposed at the market level and market-level demands are formulated and estimated. Hence, the alternative definition of structural change is used. Assumptions about utility-maximizing behavior by individual consumers are not disregarded as the model explicitly incorporates symmetry, homogeneity, and additivity conditions arising from constrained utility maximization. The symmetry condition causes structural change in the demand for one meat with respect to the price of a substitute to be reflected also as a change in the demand structure for the substitute.

Theoretical Model

A complete system of demand functions for five commodities, beef (i = 1 or B), pork (i = 2 or P), chicken (i = 3 or C), other foods (i = 4 or O) and nonfoods (i = 5 or N), is

$$(1) \quad q_i^o = q_i(y, p_1, p_2, p_3, p_4, p_5), \quad i = 1, 2, 3, 4, 5,$$

where q_i^o is per capita consumption of good i, p_i is the nominal price of good i, and y is real per capita income.[3] Multiplying the differenced logarithmic first-order Taylor series approximation of each equation by its respective budget share gives the Rotterdam specification for this system as

$$(2) \quad \omega_i d \ln q_i = \mu_i d \ln y + \sum_{j=1}^{5} \pi_{ij} d \ln p_j, \quad i = 1, 2, 3, 4, 5,$$

where ω_i is the budget share of good i, μ_i represents $\omega_i \eta_{iy}$, π_{ij} represents $\omega_i \eta_{ij}$, η_{ij} is the compensated cross-price elasticity of good i with respect to price j, and η_{iy} is the income elasticity of good i. The assumption of individual consumers' utility maximization implies that demands are additive, i.e., $\sum_{j=1}^{5} \mu_i = 1$; homogenous, i.e., $\sum_{j=1}^{5} \pi_{ij} = 0$; and symmetric, i.e., $\pi_{ij} = \pi_{ji}$. To impose these restrictions, prices are ex-

[3] The use of nominal prices with real incomes follows Theil's "Rotterdam model in absolute prices" (1975, pp. 48–49). This model is derived from compensated demand functions (as apparent from Theil 1975, p. 49) instead of the more frequently used Marshallian demand functions. Hence, real income is the appropriate argument in the demand function.

pressed as price relatives, symmetry is imposed at estimation and one equation is deleted from the system.[4] The restricted form of the Rotterdam system becomes

$$\text{(3a)}\quad \omega_i d \ln q_i = \mu_i d \ln y + \sum_{j=1}^{4} \pi_{ij} d \ln\left(\frac{p_j}{p_5}\right), \quad i = 1, 2, 3, 4,$$

or in matrix form,

$$\text{(3b)}\quad \text{diag}\ \omega\ d \ln \mathbf{q} = \mu d \ln y + \Pi d \ln \mathbf{p},$$

where the elements of the vector p are price relatives and Π is symmetric.

In this demand system, the annual supplies of beef, pork and chicken, are assumed to be fixed.[5] This assumption dictates that meat quantities should be treated as given along with the prices of other foods and nonfood items and that meat prices should be dependent variables. Accordingly, the vectors $\mathbf{p}$ and $\mathbf{q}$ in (3b) are partitioned into endogenous and exogenous components. The price relatives, $p_2 = [p_4/p_5]$, and the complementary set of meat quantities, $\mathbf{q}_1$, are exogenous; and the remaining variables, $\mathbf{p}_1$, the meat price relatives, and $\mathbf{q}_2$, the quantity of other foods, are endogenous. Solving for the endogenous variables of the system and recalling the definitions of Π and μ gives

$$\text{(4a)}\quad \text{diag}\ \omega_1 d \ln \mathbf{p}_1 = \text{diag}\ \omega_1[(-\eta_{11}^{-1}\eta_{1y}) d \ln y + \eta_{11}^{-1} d \ln \mathbf{q}_1 + (-\eta_{11}^{-1}\eta_{12}) d \ln \mathbf{p}_2],$$

$$\text{(4b)}\quad \text{diag}\ \omega_2\ d \ln \mathbf{q}_2 = \text{diag}\ \omega_2[(\eta_{2y} - \eta_{21}\eta_{11}^{-1}\eta_{1y}) d \ln \cdot y + \eta_{21}\eta_{11}^{-1} d \ln \mathbf{q}_1 + (\eta_{22} - \eta_{21}\eta_{11}^{-1}\eta_{12}) d \ln \mathbf{p}_2],$$

where the income and price elasticity matrices are partitioned to correspond to the partitioning of $\mathbf{p}$ and $\mathbf{q}$.[6] This system is similar to the original Rotterdam model except that the beef, pork, and chicken equations, represented by (4a), are in price-dependent form. The parameters to be estimated are the elements of Ω_{1y}, Ω_{2y}, Ω_{11}, Ω_{12}, Ω_{21}, and Ω_{22} in

$$\text{(5a)}\quad \text{diag}\ \omega_1 \mathbf{d} \ln \mathbf{p}_1 = \Omega_{1y} \mathbf{d} \ln y + \Omega_{11} \cdot d \ln \mathbf{q}_1 + \Omega_{12} d \ln \mathbf{p}_2,$$

$$\text{(5b)}\quad \text{diag}\ \omega_2 d \ln \mathbf{q}_2 = \Omega_{2y} d \ln y + \Omega_{21} \cdot d \ln \mathbf{q}_1 + \Omega_{22} d \ln \mathbf{p}_2.$$

The symmetry properties become $\Omega_{11} = \Omega_{11}^T$, $\Omega_{12} = -\Omega_{21}^T$, and $\Omega_{22} = \Omega_{22}^T$. The parameters in Ω can be used to find the compensated demand elasticities by using the correspondence between the structural parameters in (4a) and (4b) and the reduced form parameters in (5a) and (5b).[7] A flexibility matrix can also be derived by inverting the complete Marshallian elasticity matrix.

The model in (5a) and (5b), with its symmetry restrictions, can be estimated as

$$\text{(6a)}\quad Y = X\Theta + u \quad \text{or} \quad Y_{it} = X_{it}\Theta + u_{it}, \quad i = 1, 2, 3, 4,$$

where the subscript t designates year less 1900; Y_{it} represents $\omega_{i,t-1}\Delta \ln (p_{it}/p_{5t})$ for $i = 1, 2, 3$ and $\omega_{i,t-1}\Delta \ln q_{it}$ for $i = 4$; X_{it} is a row vector (with 14 columns) of differenced logarithms of the exogenous variables, which is constructed so that the symmetry conditions hold upon estimation; Θ is a column vector of structural parameters that correspond to the unique elements of Ω; and u_{it} are random errors. Under the assumption that $E(u_{it}, u_{jt'}) = \sigma_{ij}$ if $t = t'$, and zero otherwise, the Zellner seemingly unrelated regressions estimator (Pindyck and

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[4] The parameters in the deleted equation are derived from the parameters of the other equations through the symmetry and homogeneity conditions. The necessity of deleting one equation from the system is due to the additivity of expenditures to total income, which results in a singular contemporaneous covariance matrix. Theil (1975) points out that the choice of the deleted equation is benign. A more formal proof is available in Theil (1971).

[5] This assumption is based on the length of the gestation and production cycles for beef and pork and empirical evidence for chicken. Heien estimated the annual supply price elasticity for broilers to be .11, and Chavas and Johnson estimated that it takes about twenty quarters for broiler production to react fully to a change in the wholesale price. The appeal made here is that the simultaneous equations bias for price-dependent demand functions → 0 as the supply price elasticity → 0, *ceteris paribus*. Because the elasticity is small, the simultaneous equations bias for a price-dependent demand function is smaller than for a quantity-dependent demand function, although two recent studies of chicken demand (Thurman 1986, 1987) contradict this claim. Ideally, a simultaneous system should be modeled, but the computational burden of performing an extensive grid search (about 17,500 individual regressions) with an instrumental variables estimator is overwhelming. The maintenance of public and private inventories of other food commodities and the openness of the U.S. economy to international trade in nonfood suggest an infinitely elastic supply behavior for these commodities.

[6] Equations (4a) and (4b) represent an algebraic rearrangement of (3b) after partitioning according to its endogenous and exogenous arguments. Using the definitions established both before and after (3b) and letting $\bar{\omega}_1 = \text{diag}\ \omega_1$ and $\bar{\omega}_2 = \text{diag}\ \omega_2$ gives

$$\bar{\omega}_1 d \ln q_1 = \bar{\omega}_1\eta_{1y} d \ln y + \bar{\omega}_1\eta_{11} d \ln p_1 + \bar{\omega}_1\eta_{12} d \ln p_2$$

$$\bar{\omega}_2 d \ln q_2 = \bar{\omega}_2\eta_{2y} d \ln y + \bar{\omega}_2\eta_{21} d \ln p_1 + \bar{\omega}_2\eta_{22} d \ln p_2.$$

Solving the first of these equations for the endogenous term $d \ln p_1$ gives (4a). Substituting this expression for $d \ln p_1$ in the second equation gives (4b).

[7] This correspondence is

$$\Omega_{1y} = -\bar{\omega}_1\eta_{11}^{-1}\eta_{1y} \qquad \Omega_{11} = \bar{\omega}_1\eta_{11}^{-1} \qquad \Omega_{12} = -\bar{\omega}_1\eta_{11}^{-1}\eta_{12}$$

$$\Omega_{2y} = \bar{\omega}_2(\eta_{2y} - \eta_{21}\eta_{11}^{-1}\eta_{1y}) \qquad \Omega_{21} = \bar{\omega}_2\eta_{21}\eta_{11}^{-1} \qquad \Omega_{22} = \bar{\omega}_2(\eta_{22} - \eta_{21}\eta_{11}^{-1}\eta_{12})$$

where $\bar{\omega}_1 = \text{diag}\ \omega_1$ and $\bar{\omega}_2 = \text{diag}\ \omega_2$. Given the estimated Ω, the compensated elasticities can be computed starting with $\eta_{11} = \Omega_{11}^{-1}\bar{\omega}_1$.

Rubinfield, pp. 347–49) is consistent and asymptotically efficient.

A structural change specification is added to (6a) by fitting the regression model

$$\text{(6b)} \qquad Y_{it} = X_{it}\Theta + Z_{it}\Phi + u_{it}, \qquad i = 1, 2, 3, 4,$$

$$\text{or} \quad Y = X\Theta + Z\Phi + u,$$

where the new terms are Φ, a column vector representing the magnitudes of the structural adjustments to Θ; and Z_{it}, formed as $Z_{ikt} = \delta_{kt} \cdot X_{ijt}$, $k = 1, 2, \ldots, K$, where the δ_{kt}'s are dummy variables that depend on time, t.[8] Under the stochastic assumptions stated previously, the Zellner estimator is also appropriate for the estimation of (6b) and was used.

The usual treatment in creating column k of Z is to let δ_{kt} equal 0 if $t < \tau_k$ and equal one if $t \geq \tau_k$. The problems with this treatment are, first, the point of structural change, τ_k, must be specified a priori; and, second, the model cannot move gradually from one structural regime to the next. A more general form of structural change can be created by replacing the 0, 1 dummy variable with either a logit variable,

$$\text{(7a)} \qquad \delta_{kt} = \lambda(t; \beta_k, \tau_k) = \frac{1}{1 + e^{-\beta_k(t-\tau_k)}},$$

or an exponential variable,

$$\text{(7b)} \qquad \delta_{kt} = \gamma(t; \beta_k, \tau_k) = \begin{cases} 1 - e^{-\beta_k t} & t \geq \tau_k \\ 0 & \text{otherwise} \end{cases}.$$

These models generalize the simple dummy variable model in that in both cases, as β_k gets large, the functions take values of zero to the left of τ_k and take values of one to the right of τ_k. Additionally, gradual structural change can be modeled when β_k takes small values.

When these gradually shifting dummy variables are used in (6b), the total structural change, $\Sigma_k Z_{ikt}\Phi_k$ with $Z_{ikt} = \delta_k(t; \beta_k, \tau_k)X_{ijt}$, can be decomposed. First, information about the timing, rate, and magnitude of each of the K changes is conveyed by τ_k, β_k, and Φ_k, respectively. The origin of structural change is the conditioning variable in column j of X, whose parameter appears to change. The object of structural change is the dependent variable in equation i, where the structural change occurs. The time path of the effect of independent variable j on dependent variable i is Θ_j plus the sum of all $\Phi_k\delta_k(t; \beta_k, \tau_k)$ that interacted with regressor j.[9]

This model was estimated using annual data covering a period from 1950 through 1985. Available in *Food Consumption, Prices, and Expenditures* (USDA) were data on per capita beef, pork, and chicken consumption; retail weight equivalent of total food consumption; personal consumption expenditures for food and for all goods and services; total U.S. population; and consumer price indices for all items and for food only. Annual average retail beef, pork, and chicken prices are available in *Livestock and Poultry Outlook and Situation Report* (USDA).[10] Per capita personal consumption expenditure deflated by the consumer price index was used to represent real consumer income.[11] A proxy for other food consumption was derived by subtracting beef, pork, and chicken consumption from the retail weight equivalent of total food consumption. The average price of other food consumed was computed by dividing expenditures on other food by the quantity of other food. The CPI excluding food items was used as the price for nonfood and was available in the *Statistical Abstract of the United States.*

Estimation Procedure and Results

Estimation of (6b) requires that K, the column dimension of Z, or equivalently, the total number of structural changes, be selected. For

[8] Each of the K columns of Z corresponds to a structural change. An alternative treatment is to define $Z_{ijt} = \delta_{jt}X_{ijt}$, $j = 1, 2, \ldots, 14$, so that Z has 14 columns, each corresponding to the same column of X. This alternative treatment allows each element of Θ to display only one structural change. The specification selected is more general because it allows each element of Θ to display several changes.

[9] In other words,

$$\partial Y_{it}/\partial X_{ijt} = \Theta_j + \sum_{k=1}^{K} \Phi_k(\partial Z_{ikt}/\partial X_{ijt})$$

where

$$\partial Z_{ikt}/\partial X_{ijt} = \begin{cases} 0 \text{ if commodity } j \text{ is not the source of structural change } k, \text{ or} \\ \delta_k(t; \beta_k, \tau_k) \text{ otherwise.} \end{cases}$$

[10] These prices were weighted average prices of retail cuts for beef and pork and four-region-average prices for young chickens.

[11] The income constraint in a constrained utility-maximization problem states that total expenditures equals income. The choice of income variables is therefore between income, represented by per capita disposable personal income, and expenditures, represented by per capita personal consumption expenditures. The additivity condition requires that the data used to represent income should be such that additivity applies. Additivity was maintained by using per capita personal consumption expenditures to represent income because per capita personal consumption expenditures was equal to total expenditures on the commodities in the system.

each individual change, the type of the change, either logit or exponential, must be selected, and the corresponding parameters, τ_k and β_k, must be estimated. The model is nonlinear in the τ_k's and the β_k's but is linear in the parameters of the original structure, Θ, and the magnitude parameters of the structural changes, Φ.

Stepwise likelihood maximization was used to estimate (6b). One advantage of this procedure over direct estimation was that it was more parsimonious in the number of parameters to be estimated. All possible structural change specifications did not need to be included in the model at the outset, because the procedure searched for the most significant specifications, which were then added to the model. This procedure also allowed specifications in which an equation could exhibit multiple changes in its response to a single explanatory variable. Another advantage of this procedure over direct estimation is that direct estimation could lead to a singular regressor matrix and the failure of the algorithm to converge.[12]

In general terms, the search was conducted in steps. At each step, the most significant specification of structural change was found and added to the model as a column of Z, and a test was conducted to determine if the added column explained a significant amount of variation in the dependent variable. In detail, the algorithm is as follows:

(1) Fit (6a), which assumes no structural change, i.e., set $K = 0$. Also set $k = 0$.

(2) Increment k by 1 and consider adding $\delta_{kt}X_{ijt}$ to the model. Select the shape of $\delta_k(t)$ and the origin (i.e., column j of X) of the structural change by searching for the maximized value of the likelihood function where the likelihood values are computed from the iterative Zellner estimator of (6b). Consider all possible combinations of (*a*) structural change types, either logit (7a), or exponential (7b); (*b*) structural change origins, i.e., the $j = 1, 2, \ldots, 14$ columns of X; and (*c*) structural change shapes, i.e., $\delta_k(t; \beta_k, \tau_k)$ using values of τ_k that correspond to the years from 1955.5 to 1980.5 by 1 and β_k from 2^{-6} to 2^4 by powers of 2.

(3) Refine the estimates of τ_k and β_k by searching for the maximized value of the likelihood function using smaller steps in a smaller neighborhood around the maximizing values in 2. Determine β_k to the nearest hundredth and τ_k to the nearest tenth.

(4) Test the statistical significance of the increase in the likelihood function by using the likelihood ratio test with three degrees of freedom (one each for β_k, τ_k, and Φ_k). If significant, column k of Z becomes the $Z_{ikt} = \delta_{kt}X_{ijt}$, corresponding to the maximum likelihood function value. Set K to k and search for an additional structural change by repeating 2 and 3. If the increase is not significant, terminate the search.

The likelihood function was evaluated under approximately 17,500 structural change specifications and three distinct structural shifts were detected.

Table 1 summarizes the estimation path. This table is ordered by steps, k, where at each step Z contains k columns. The equation-origin combination indicates which variable (origin) appears to exert a new influence on a commodity (equation). The estimated β_k and τ_k determine the shape of the structural change when substituted into the indicated logit or exponential functions. Two times the natural logarithm of the maximized value of the likelihood function is denoted as $2 \ln L^*$, where the asterisk indicates the maximum of the two functional forms of structural change. The likelihood-ratio test statistic is the difference between successive values of $2 \ln L^*$. These differences are distributed as chi-square random variables with three degrees of freedom. The final column shows the probability of a larger chi square under the null hypothesis of no structural change at step k.

The base model, (6a), has a (2 times the logarithm of) likelihood value of 360.3638. Adding the first structural change specification increased the maximized likelihood function value giving $2 \ln L^*$ of 380.8171. Comparison of the logit and the exponential forms of change reveals that creating a variable from the exponential function times the column of X containing $\Delta \ln QP_t$ increased the maximized value of the likelihood function more than did the same product using a logit function. The likelihood ratio test for the null hypothesis $\Phi_1 = 0$, and/or τ_1 and β_1 such that $\delta_{1t} = 0$ for all t gave the probability of a larger chi square of .0001 so that the null hypothesis was rejected. The first column of Z was created as $\Delta \ln QP_t \times [1 - e^{-.22(t-55.0)}]$.

In step 2, the likelihood function increase is

[12] Estimation of τ_k and β_k creates the possibility of generating a singular regressor matrix when (1) $\beta_k \to 0$, or when (2) $\tau_k \to \pm\infty$, resulting in either $\delta_k(t) = 1$ or $\delta_k(t) = 0$ for all t.

Table 1. Stepwise Likelihood Maximization Summary

Step (k)	Structural Change			Estimated		2 ln L^*	Pr > χ^2
	Equation	Origin	Form	β_k	τ_k		
0	Base model—no structural change					360.3638	
1	Pork	$\Delta \ln QP_t$	Logit	1.36	57.8	380.4726	
			Exp	.22	55.0	380.8171*	.0001
2	Beef	$\Delta \ln QC_t$	Logit	32.0	69.5	389.2389*	.0381
	Chicken	$\Delta \ln QB_t$	Exp	32.0	69.5	389.2389*	.0381
3	Beef	$\Delta \ln QC_t$	Logit	.97	75.8	397.8155*	.0355
	Chicken	$\Delta \ln QB_t$	Exp	.61	74.4	397.7747	
4	Other food	$\Delta \ln Y_t$	Logit	32.0	63.5	402.5995*	.1883
			Exp	32.0	63.5	402.5995*	.1883

greatest when either the logit or exponential functions are multiplied by the column of X that contains the symmetric effects of beef production on the chicken price and chicken production on the beef price. Hence, the two-equation-origin combinations represent a change in a parameter that is restricted by symmetry. Table 1 indicates that the likelihood function is maximized at $\beta_2 = 32.0$, where both the logit and the exponential functions behave computationally as 0, 1 dummy variables. This structural change was significantly nonzero at beyond the 5% level.

Step 3 in table 1 indicates another change in the symmetric relationships between beef prices and chicken production and between chicken prices and beef production. The estimated timing, τ_3, was later and the estimated rate of adjustment, β_3, was slower than the corresponding parameters estimated in step 2. The logit dummy variable specification resulted in a larger maximized likelihood function value than did the exponential model. This effect was significantly nonzero at beyond the 5% level, so it was used as the third column of Z.

At the fourth step, the likelihood function is found to have the greatest increase by including a change in the way other food consumption reacts to income. However, this effect is not significantly different from zero at beyond the 5% significance level. Thus, the search terminates and the number of regressors in Z is 3.

Table 2 shows estimates of the original demand-structure parameters, Θ, and of the

Table 2. Estimated Parameters for the Mixed Rotterdam Model

Parameter	Estimate	Standard Error	T-Ratio	Significance
$\Theta_1 = \Omega_{BB}$	−2.9984	.3971	−7.5513	.0000
$\Theta_2 = \Omega_{BP}$	−.7973	.2135	−3.7344	.0003
$\Theta_3 = \Omega_{BC}$	−.1980	.1022	−1.9368	.0548
$\Theta_4 = \Omega_{BO}$	.2807	.2996	.9370	.3504
$\Theta_5 = \Omega_{BY}$	1.3205	.7835	1.6854	.0941
$\Theta_6 = \Omega_{PP}$	−3.7444	.3160	−11.8512	.0000
$\Theta_7 = \Omega_{PC}$	−.2685	.0694	−3.8672	.0002
$\Theta_8 = \Omega_{PO}$	.2093	.2081	1.0055	.3164
$\Theta_9 = \Omega_{PY}$	.3008	.4771	.6305	.5294
$\Theta_{10} = \Omega_{CC}$	− .7253	.0892	−8.1341	.0000
$\Theta_{11} = \Omega_{CO}$	.2265	.1681	1.3472	.1801
$\Theta_{12} = \Omega_{CY}$	.2619	.2061	1.2710	.2058
$\Theta_{13} = \Omega_{OO}$	−1.6550	.6891	−2.4017	.0176
$\Theta_{14} = \Omega_{OY}$	1.1667	.6547	1.7820	.0769
$\Phi_1 = \Delta\Omega_{PP}$	1.8937	.3120	6.0700	.0000
$\Phi_2 = \Delta\Omega_{BC}$	−1.2537	.2667	−4.7007	.0000
$\Phi_3 = \Delta\Omega_{BC}$	1.1550	.3438	3.3590	.0010
Weighted system R^2	.8177			

structural-change magnitude parameters, Φ. These parameter estimates are identified both by their positions in the Θ and Φ vectors of (6b) and by their positions in the Ω matrix of (5a) and (5b). The structural-change magnitude parameter estimates—i.e., the Φ's—are highly significant, while the significance levels of the other parameters are mixed. The results show that the production of each meat significantly and negatively influences the price of that meat as well as having strong depressing effects on the prices of other meats. The effects of income and the other-food price on meat prices are not highly significant, but because these parameters estimate neither elasticities nor flexibilities, there are no strong a priori expectations about their signs. This table also shows that the other-food price has a strong negative effect on other-food consumption and that income has the expected positive impact on other-food consumption.

The estimated rate and timing parameters, β_k and τ_k from table 1, and the estimated magnitude parameters, Φ from table 2, can be combined with the estimated base-structure parameters, Θ, also from table 2, to derive the time-varying reduced-form parameters shown in figure 2. Figure 2a shows that adjustment in the direct pork price parameter began in 1955 and followed an exponentially damped path. Concerns about the current implications of this phenomenon can be dismissed because the adjustment was completed by 1960.

Figure 2b reflects the two adjustments in the beef-chicken cross-commodity parameter. This parameter first displays a discrete change between 1969 and 1970, as detected in step 2 of table 1. Then, as detected in step 3 of table 1, the parameter starts back toward its original level, but it takes most of the 1970s to reach that level. The return of the parameter to its original level or, alternatively, the near-equality but opposite sign of the two magnitude parameters associated with these changes, gives rise to the hypothesis that the system is returning or has returned to its 1960s, or long-run, structure. The rejection of the hypothesis, $\Phi_2 = -\Phi_3$, would discredit this notion. Fitting a model restricted by $\Phi_2 = -\Phi_3$ and applying the likelihood ratio test results in a chi-square-test statistic with a probability of a larger value of .2515. The hypothesis cannot be rejected.

The conclusion, based on these empirical results, is that meat demands display considerable stability. Any or all of the fourteen base

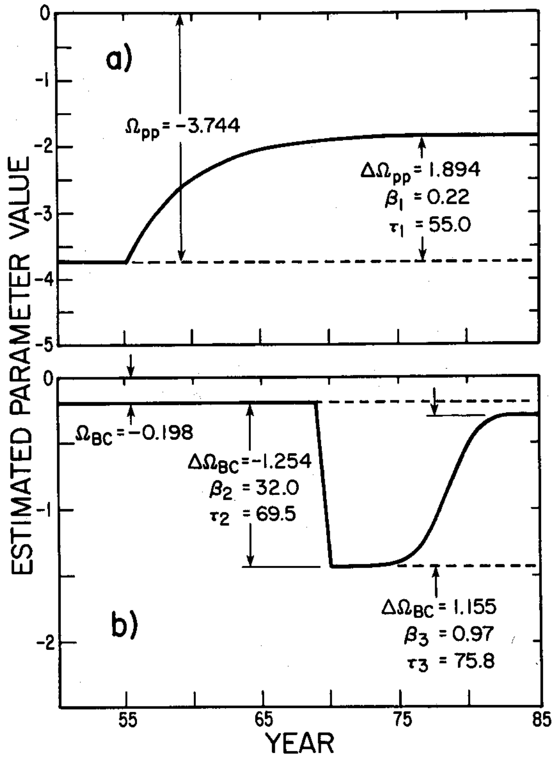

Figure 2. Behavior of time-varying model parameters over sample period, (*a*) the direct pork price parameter, and (*b*) the symmetry-constrained beef-chicken parameter

structure parameters were allowed to adjust any number of times to capture structural change in any of the meat, other-food or nonfood demands, but only two of the parameters displayed a total of three adjustments. Though the empirical model is not suitable for determining the exact cause of these adjustments, speculation about possible causes is nonetheless of interest. Possible reasons for the detected adjustments are (*a*) specification error, in that the algebraic form of the Rotterdam demand model does not coincide with the algebraic form of the underlying demands; (*b*) changes in consumers' preferences; or (*c*) specification error, in that influences other than those included in the model also influence per capita consumption.

One way to examine the possibility of algebraic misspecification is to examine the assumed constancy of the parameters π_{ij} and μ_i in the Rotterdam model. If the constancy of π_{ij} and μ_i does not accurately represent the underlying demands, then a more general specification is

(8) $\pi_{ij} = \pi_{ij}(P, Y, \gamma)$ and $\mu_i = \mu_i(P, Y, \gamma)$,

where P is a vector of prices, Y is income, and γ is a vector of unobservable utility function parameters. Although the explicit functional forms of (8) are unknown, these functions can be differentiated and the partial derivatives estimated with regression analysis to determine if, over the sample period, the $d\pi_{ij}$ and $d\mu_i$ correspond to changes in any of the observable explanatory variables. Obviously, if a parameter does not change, then the $d\pi_{ij}$ or $d\mu_i$ are zero at each point in the sample period and there is nothing to be explained by the observable explanatory variables.

To detect algebraic specification error, period-to-period changes in the pork parameter, and period-to-period changes in the beef-chicken parameter were each regressed against period-to-period changes in the exogenous variables, i.e., quantities of beef, pork, chicken; other-food and nonfood prices; and per capita income. The resulting regression F-statistics of 2.28 and .66, respectively, were significant at the .059 and .682 levels. The hypothesis of no relationship between changes in the model's beef-chicken parameter and changes in the economic variables could not be rejected, but it is tempting to reject a similar hypothesis for the pork parameter.

Upon closer inspection, changes in the pork parameter were found to be significantly and negatively correlated with changes in the other-food price, the nonfood price, and nominal per capita income, individually. However, when all of the explanatory variables were put in a single regression, no individual variables were significant because of multicollinearity. Taken together, the absence of robustness in the relationship between the pork parameter changes and changes in the explanatory variables, the lack of relationship between changes in the explanatory variables and changes in the beef-chicken parameter, and the stability of the other model parameters tend to discredit the notion of serious algebraic misspecification of the model.

The second possible source of the estimated parameter adjustments is a change in the structure of consumers' preferences. Although the initial change in the beef-chicken parameter is consistent with increased substitutability between beef and chicken because of preference changes, the fact that this parameter returns to levels not significantly different from its original level is not consistent with a permanent change in tastes. The overall stability of the remainder of the parameter structure also suggests that preferences are stable over the sample period.

Permanent changes in consumers' preferences may have been responsible for the adjustment in the pork parameter, but an equally plausible explanation links the behavior of the parameter to the characteristics of the product. The barrow-and-gilt slaughter weight (*Agricultural Statistics,* USDA), trended to a minimum of 222 pounds in 1957 and trended steadily upward after that time. Regressing these slaughter weights on the pork parameter results in an R^2 of .66 which is significant beyond the .001 level. It is possible that the observed adjustment in the pork parameter reflects either a change in consumers' tastes for pork, or a change in the commodity pork as, over time, meatier hogs were slaughtered.

The change in the chicken-beef substitution relationship is left to be explained. The timing of the change points to specification error caused by the exclusion of important influences from the model. The apparent change, which occurred in the 1970s, may have been caused by macroeconomic shocks, such as the wage and price controls imposed during the Nixon administration, and incorrectly perceived inflation. Wage and price controls obviously violate the assumptions of the model by causing shortages. If the accelerating inflation of the 1970s was incorrectly perceived, consumers' assessments of real prices may have been erroneous. Furthermore, the high inflation rates of the 1970s may have distorted consumption-savings decisions which are not considered in the model.

Other factors that may possibly have affected meat demands during this period were price controls on meats, which caused shortages and product substitutions for nonprice reasons; overreaction to new information about cholesterol in the diet; and the increased labor force participation by women with the subsequent revaluation of time spent in meal preparation. The transitory nature of these shocks is consistent with the observed transitory disruption of the beef-chicken substitution relationship. Ultimately, the parameter structure of the 1980s is like the parameter structure of the 1960s.

The long-term stability of the reduced-form parameters, however, does not necessarily im-

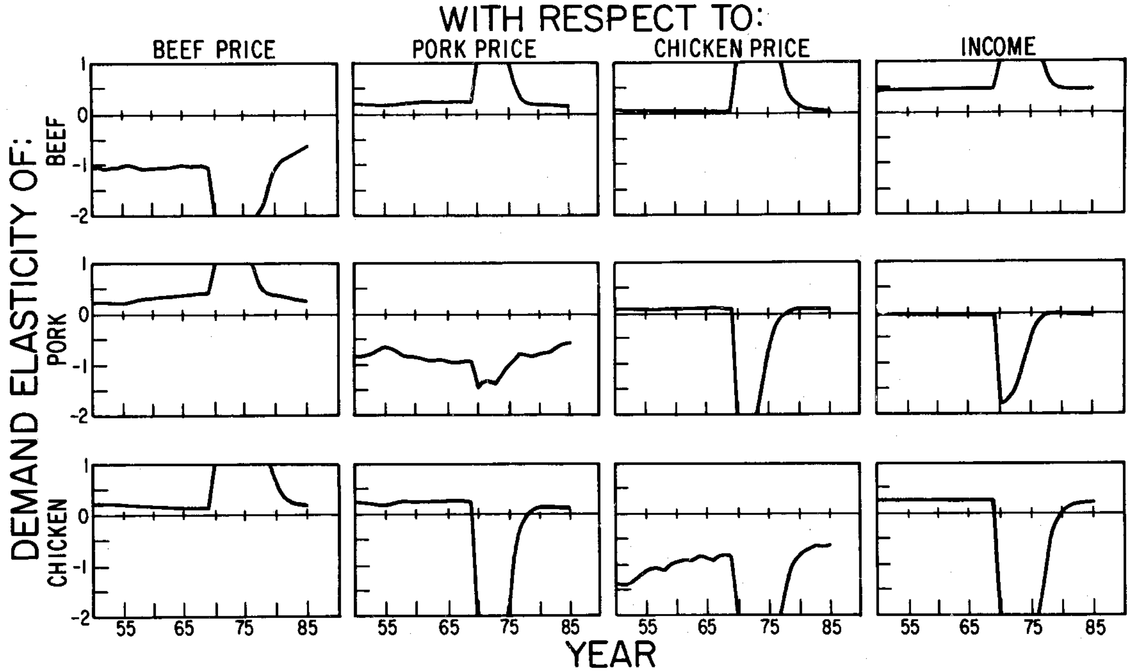

Figure 3. Derived behavior of meat demand direct-price, cross-price, and income elasticities over time

ply stability in elasticity structure of the markets. The time-varying parameter matrix $\Omega(t)$ was used to derive the Slutsky elasticity matrix. The most important partitions of this matrix are

(9a) $$\eta(t)_{11} = [(\text{diag } \omega(t)_1)^{-1}\Omega(t)_{11}]^{-1}$$

and

(9b) $$\eta(t)_{1y} = -\ \eta(t)_{11}(\text{diag } \omega(t)_1)^{-1}\Omega(t)_{1y},$$

where $\eta(t)_{11}$ is a time-varying matrix of direct- and cross-price compensated elasticities for the meats and $\eta(t)_{1y}$ is a time-varying vector of income elasticities of meat demand. The time-varying Marshallian elasticity matrix for the meat demands was derived from the compensated elasticity matrix, and the key elements of the Marshallian matrix are plotted over time in figure 3.[13] The vertical scales of the plots run from −2 to 1 so as to include a range of values appropriate for direct-price, cross-price, and income elasticities. Plotting these elasticities over time, shown on the horizontal axis, allows visual inspection of their temporal behavior.

The most prominent feature of figure 3 is that the adjustment in the beef-chicken cross-price relationship during the 1970s severely disrupted the apparent elasticity structure of the markets. All twelve elasticities show this disruption, and in some plots the elasticities move outside the range of the vertical axes and take unbelievably large (or small) values. Apart from this disruption, however, the markets appear to have a fairly stable elasticity structure. All of the cross-price and income elasticities exhibit nearly constant behavior before the disruption; and after the disruption, these elasticities restabilize at nearly the same values displayed before the disruption. The direct beef and pork price elasticities were fairly stable before the disruption but appear to be converging at less elastic levels after the disruption. Finally, the direct chicken price elasticity appears to have been trending toward less elasticity before the disruption, and the trend appears to be continuing after the disruption.

[13] The conversion between Marshallian and Slutsky, or compensated, elasticities is (Johnson, Hassan, and Green, p. 32)

$$\eta^m_{ij} = \eta^s_{ij} - \omega_j\eta_{iy},$$

where η^m_{ij} and η^s_{ij} are, respectively, Marshallian and Slutsky elasticities. The discrepancies between the Marshallian and the Slutsky elasticities for meat demands are small because of the small budget shares and small income elasticities for the meats.

Table 3. Comparison of Elasticities and Flexibilities, 1960s versus 1985

Commodity	Price and Income Elasticities					Flexibilities[a]		
	Beef	Pork	Chicken	Other Food	Income	Beef	Pork	Chicken
				Average for the 1960s				
Beef	−1.041	.218	.029	−.000	.438	−1.069	−.273	−.071
Pork	.369	−.914	.097	.043	−.056	−.460	−1.243	−.156
Chicken	.138	.273	−.863	.242	.262	−.268	−.400	−1.184
				1985				
Beef	−.659	.136	.040	−.004	.435	−1.725	−.439	−.160
Pork	.255	−.584	.069	.055	−.054	−.812	−1.960	−.267
Chicken	.170	.156	−.602	.238	.202	−.626	−.581	−1.720

[a] Flexibilities are defined as the percent change in the price of the commodity shown on the left-hand margin per 1% change in quantity supplied of the commodity listed across the top.

Conclusions

Several conclusions can be drawn from this research. First, significant changes in the demand system parameters were detected, and these changes were consistent with increased substitutability between beef and chicken. However, the timing and transitory nature of these changes does not support the contention of a permanent change in consumers' meat consumption preferences. Thus, the conclusion is that the departure from long-term meat consumption trends, depicted in figure 1, is most likely the result of changing supply conditions interacting with stable meat demands.

Second, corresponding to the detected model parameter changes, the meat demand elasticity structure appeared to change substantially in the 1970s, but in the 1980s it has restabilized. The evidence indicates that the 1970s structure was an aberration and that the meat markets have since returned to an elasticity structure that is not very different from that displayed in the 1960s. It appears, however, that direct price elasticities are smaller and may decrease still further in the future if the trends shown in figure 3 continue. Barring severe macroeconomic disturbances, such as those in the 1970s, the prediction can be made that meat demands will remain stable.

However, the stability of the meat-demand-elasticity structure does not mean that meat market participants can operate under the rules used in the 1960s. Table 3, which contains some key elasticities and flexibilities for the 1960s and for 1985, can be used to demonstrate how these rules have changed.[14] This table shows that all of these flexibilities have increased in absolute value. These larger flexibilities mean that the percentage price changes caused by a 1% change in quantity supplied are now larger than they were formerly. If (*a*) the flexibility point estimates are given, (*b*) price variability is generated entirely by supply variability, and (*c*) the variability of supply is constant both within and outside of the sample period are assumed, then the reduced flexibilities imply that the standard deviations of the percentage price changes for beef, pork, and chicken are now about one and one-half times the corresponding standard deviations for the 1960s.[15] Furthermore, because prices are now two to three times higher than they were in the 1960s, equivalent supply percentage changes now will cause greater absolute price variation. Under the assumptions listed, the standard deviations of one-period-ahead prices for beef, pork, and chicken, respectively, are now 4.5, 4.0, and 2.8 times their 1960s values. Hence, the entire meat industry, from producers to retailers, may want to examine their exposure to price risk caused by routine supply adjustments and perhaps revise risk management programs. Furthermore, a comparison of the flexibilities for the 1960s and for 1985 indi-

[14] The flexibilities are from the meat partition of the complete price flexibility matrix which was formed by inverting the complete Marshallian demand elasticity matrix.

[15] Mathematically, $d \ln p_i = f_i d \ln q$ where f_i is a vector of flexibilities for commodity i and $d \ln q$ is a stochastic vector of percentage changes in quantities supplied. Thus,

$$V(d \ln p_i) = f_i V(d \ln q) f_i'.$$

Assuming

$$d \ln p_{it} = (p_{it} - p_{i,t-1})/p_{i,t-1}$$

with $p_{i,t-1}$ given, gives

$$V(p_{it}) = p_{i,t-1}^2 f_i V(d \ln q) f_i'.$$

cates that flexibility-based price forecasting models need revision.

The final implication of the changing meat-market elasticity structure has to do with retail meat pricing. This implication rests on assumptions (*a*) that grocery stores are not perfectly competitive firms because of shoppers' time and travel costs to shop several stores, and (*b*) that the less elastic per-capita market-level demands reflect consumer behavior so that individual consumers' demands for meats have also become less elastic. Faced with less elastic demands and with all else constant, grocery stores would seek greater margins in the meat department and would accordingly raise meat prices. This prediction requires caution, however, because grocers' markets do not correspond to the aggregate market. Through time, additional grocery stores and competitive reactions at the grocer's market level may mitigate the price-enhancing impact of more inelastic meat demands. Whatever the case, the implications of less elastic demand for retail meat pricing remains an interesting research issue.

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