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Summary

Choice of functional form and structural change specification may each influence demand parameter estimates. Joint non-nested testing of both is applied for the linearized almost ideal and Rotterdam models with and without structural change, incorporated by the gradual transition approach. The test of model choice without structural change is inconclusive; the test of models with structural change shows that, for Canadian meat consumption, the gradual transition almost ideal model is preferred over the gradual transition Rotterdam model. The effect of functional form on demand elasticities for meats is relatively minor but the structural change provision substantially affects these estimates.

Keywords: meat demand; joint test; functional form; structural change

1. Introduction

Applied economists have focussed much attention on the demand for meat, in order to assess the basis of the appreciable changes in consumption patterns and budget allocations for individual meats that have been evident over the past two decades. Recently, two related topics have received increasing attention in the effort to understand how consumers allocate expenditure amongst meats. One focus of attention is the choice of functional form for the demand system. Another is the issue of whether structural change has occurred in the demand for meat. This paper is concerned with a joint test of functional form and structural specification appropriate to Canadian meat consumption from 1967 to 1992.

Several studies have provided methods to test separately the appropriateness of functional form and structural specifications for demand systems. Deaton (1978) first applied a non-nested test of functional form to compare demand systems which have the same dependent variables. His method can be used to compare, for example, the almost ideal with the indirect translog system since for both, budget share is the dependent variable. Barten (1993) developed a compound model to compare demand systems with identical right hand side terms. His procedure was used to choose amongst the Rotterdam, linearized almost ideal, and two related models.

Two models that are frequently used in analysis of the demand for foods are selected as the functional forms for assessment in this study. These models, the almost ideal demand system, applied in its linear approximate version, and the Rotterdam model, are similar in many respects, having identical data requirements and being equally parsimonious with respect to the number of parameters. However, the assumptions used to parameterize the two systems have rather different behavioural implications. In particular, the Rotterdam model assumes constancy of marginal expenditure shares and Slutsky terms while the almost ideal model assumes that these are functions of budget shares. Economic theory does not provide a basis for choosing *ex ante* between these two models; however, empirical studies have found different results from them.[1] Lacking *a priori* guidance from economic theory, the choice between these two models must essentially be made on the basis of empirical results.

There has been much debate regarding the existence and basis of structural change in demand for meat in recent years, sparked by the considerable change in consumption patterns for meats. The existence of structural change has been queried, based on non-parametric assessments of relatively few discrepancies in the consistency of consumers' choices over time (Chalfant and Alston, 1988; Alston and Chalfant, 1991). However, the power of these tests is not clear and parametric assessments of consumer demand for meats typically reject the null hypothesis of no structural change. Such studies, based on US data, include Thurman (1987), Moschini and Meilke (1989), Dahlgran (1988) and Choi and Sosin (1990). Studies based on Canadian data include Atkins, Kerr and McGivern (1989), Chen and Veeman (1991) and Reynolds and Goddard (1991). The conclusion of structural change has been queried by other researchers. For example, Eales and Unnevehr (1993) have suggested that supply-side shocks manifested through endogenous meat

[1] For example Alston and Chalfant (1991) report contradictory conclusions regarding structural change in comparing these two models with Canadian meat consumption data. Lee, Brown and Seale (1994), using data on Taiwanese consumption patterns, found questionable income elasticity estimates from the Rotterdam model and concluded that the almost ideal model better explained expenditure behaviour.

prices may account for apparent structural change in US meat demand. This argument is not, however, applicable to Canadian meat demand. Canada's proximity to and close market links with the US market for beef and pork exposes Canadian consumers to market prices that are determined in the much larger US market for these meats. Further, the Canadian supply-management program for chicken has involved regulatory price-setting for this product. Thus, in modelling Canadian meat demand, Canadian meat prices are not expected to be endogenous. Overall, considerable attention continues to be paid by applied economists to the issue of structural change in the demand for meat.

2. Testing for Functional Form and Structural Specifications

Two major methods have been used to model structural change in demand. One introduces a time or demographic dummy variable and assumes that the change in consumption patterns occurs abruptly. The other introduces a time transition function that allows a gradual shift from one regime to the next. This approach is applied in the paper. Since the process of testing the choice of specific models is essentially a test of fitness to a particular data set, in situations where structural change specification is also an issue, as with meats, this assessment should consider structural specification together with model choice. Ignoring structural change in testing appropriate functional form may lead to invalid test statistics due to the inability to distinguish the effects of functional and structural misspecification. To our knowledge, previous papers that have focussed on misspecification testing, as in the procedures of McGuirk *et al* (1995), have not directly dealt jointly with model choice and structural change. We adapt the compound model of Barten (1993) to test directly and jointly both functional and structural specifications using the linear almost ideal and Rotterdam models, with and without structural change, for this purpose.

3. The Gradual Switching Specification of the Almost Ideal and Rotterdam Demand Models

The Rotterdam model, ignoring the issue of structural change, takes the form:

$$w_i d\ln q_i = \beta_i DQ^* + \sum_j \gamma_{ij} d\ln p_j \qquad (1)$$

where: $(i,j)=1,...,k$ index the goods; p_j is the nominal price of good j; q_i is the quantity of good i; and $DQ^* = \Sigma_i w_i\, dlnq_i$ for $w_i = p_i q_i / x$ and $x = \Sigma_i p_i q_i$. Taking the logarithmic differential of the budget equation gives : $dlnx = \Sigma_j w_j\, dlnp_j + \Sigma_j w_j\, dlnq_j$. This is rewritten as: $\Sigma_j w_j\, dlnq_j = dlnx - \Sigma_j w_j\, dlnp_j$ and substituted into (1), giving the following version of the Rotterdam model:

$$w_i dlnq_i = \Sigma_j \gamma_{ij} dlnp_j + \beta_i (dlnx - \Sigma_j w_j dlnp_j) \qquad (2)$$

The constraints of demand theory can be directly applied to the Rotterdam parameters. These are, for adding up, $\Sigma_i \beta_i = 1$, $\Sigma_i \gamma_{ij} = 0$; for homogeneity, $\Sigma_j \gamma_{ij} = 0$; and for Slutsky symmetry, $\gamma_{ij} = \gamma_{ji.}$

The almost ideal demand model without structural change is expressed as:

$$w_i = \alpha_i + \sum_j \delta_{ij} lnp_j + \rho_i \ln(x/P) \qquad (3)$$

where P is a nonlinear price index. We follow the commonly-used procedure approximating P by Stone's geometric index, $ln(P) = \Sigma_j w_j lnp_j$. The first difference of Stone's index can be decomposed into three components:

$$dlnP = \Sigma_j w_j dlnp_j + \Sigma_j dw_j lnp_j - \Sigma_j dw_j dlnp_j \qquad (4)$$

For time series data the last two terms are usually very small (Alston and Chalfant, 1993) and typically ignored by analysts as negligible (Barten, 1993; Lee *et al*., 1994). Substituting the first term of $dlnP$ from (4) into the first differenced form of the almost ideal model yields:

$$dw_i = \Sigma_j \delta_{ij} dlnp_j + \rho_i (dlnx - \Sigma_j w_j dlnp_j) \qquad (5)$$

The theoretical properties of adding up, homogeneity, and symmetry in this system imply, respectively, that: $\Sigma_i \rho_i = 0$, $\Sigma_i \delta_{ij} = 0$; $\Sigma_j \delta_{ji} = 0$; and $\delta_{ij} = \delta_{ji.}$

Comparison of the right-hand sides of Equations 2 and 5 indicates their similarity. Their left-hand sides differ, but are related. Barten (1993) shows that this can clearly be demonstrated by taking the logarithmic differential of the budget share, which gives $dw_i = w_i dlnq_i + w_i dlnp_i - w_i dlnx$ and replacing $w_i dlnq_i$ as the right hand side of (2) yielding: $\rho_i = \beta_i - w_i$ and $\delta_{ij} = \gamma_{ij} - w_i w_j - \theta_{ij} w_i$ where θ_{ij} is the Kronecker delta, equal to unity if $i=j$ and 0 otherwise. Since w_i varies, taking ρ and δ to be constant (as in the almost ideal model) has different implications than taking γ and β to be constant (as in the Rotterdam model). The two systems are different but comparable (Barten, 1993: 145).

Under the hypothesis of absence of structural change and assuming the almost ideal (or, alternatively, the Rotterdam) model provides a satisfactory approximation of the true demand system, the set of parameters $\{\rho_i, \delta_{ij}\}$ (or, alternatively $\{\beta_i, \gamma_{ij}\}$) represent the underlying utility maximization process. Structural change can be characterized by allowing this set of parameters to change over time. If we assume a time path, h_t, that affects all equations simultaneously, the first differenced almost ideal model of Equation 5 can be reparameterized as:

$$dw_{it} = \tau_i dh_t + \Sigma_j(\delta_{ij} + \pi_{ij} h_t) dlnp_{jt} + (\rho_i + \pi_i h_t)(dlnx_t - \Sigma_j w_{jt} dlnp_{jt}) + u_{it} \qquad (6)$$

The properties of adding up, homogeneity and symmetry require, in addition to those reported for (5), the following parametric restrictions: $\Sigma_i \tau_i = 0$, $\Sigma_i \pi_{ij} = 0$, $\Sigma_i \pi_i = 0$; $\Sigma_j \pi_{ji} = 0$; $\pi_{ij} = \pi_{ji}$.

Including seasonal dummy variables, the gradual switching almost ideal demand becomes:

$$\begin{aligned} dw_{it} = {} & \tau_i dh_t + \Sigma_j[\delta_{ij} dlnp_{jt} + \pi_{ij}(h_t dlnp_{jt})] + \rho_i(dlnx_t - \Sigma_j w_{jt} dlnp_{jt}) + \pi_i[h_t(dlnx_t - \Sigma_j w_{jt} dlnp_{jt})] \\ & + \Sigma_k[\mu_{ik} dD_k + \tau_{ik} d(h_t D_k)] + u_{it} \end{aligned} \qquad (7)$$

where D_i = seasonal dummies and h_t = a transition function expressing the time path from one regime to the other, defined as:

$h_t = 0$ $\qquad$ *for* $t=1,..., t_1$

$h_t = (t-t_1)/(t_2-t_1)$ $\qquad$ *for* $t= t_1+1, ..., t_2-1$

$h_t =1$ $\qquad$ *for* $t=t_2,...,T$

where t_1 = the end point of the first regime and t_2 = the start point of the second regime. The conditions $t_1 \leq T-m/(n-1)$, $t_2 \geq m/(n-1)$, $t_1 < t_2$ are applied to guarantee sufficient observations in each regime for model estimation where n = the number of equations in the system and m = the total number of free parameters to be estimated.

The advantage of the transition function is that it allows a gradual shift from one regime to the next. Whether dynamic demand patterns are the result of changes in consumer tastes and preferences or reflective of changes in demographic structure, as has been suggested by analysts and observers, it seems reasonable to allow for gradual changes in demand. Moschini and Meilke obtained maximum likelihood estimates of t_1 and t_2 by searching the likelihood function over the range of interest of (t_1 t_2) and found that t_1 =1975-IV and t_2 =1976-III maximized the likelihood functions for US meat consumption data.

Since, as shown above, the Rotterdam model has same right hand side as the first differenced almost ideal model, the gradual switching Rotterdam model can similarly be written as:

$$w_{it}dlnq_{it}=\alpha_i dh_t+\Sigma_j[\gamma_{ij}dlnp_{jt}+\phi_{ij}(h_t dlnp_{jt})]+\beta_i(dlnx_t-\Sigma_j w_{jt}dlnp_{jt})+\phi_i[h_t(dlnx_t-\Sigma_j w_{jt}dlnp_{jt})] \\ +\Sigma_k[\omega_{ik}dD_k+\alpha_{ik}d(h_t D_k)]+e_{it} \qquad (8)$$

The properties of adding up, homogeneity and symmetry require, in addition to those specified for (2), the following restrictions on the parameters: $\Sigma_i\alpha_i=0$, $\Sigma_i\phi_i=0$, $\Sigma_i\phi_{ij}=0$; $\Sigma_j\phi_{ij}=0$; $\phi_{ij}=\phi_{ji}$.

Equations (7) and (8) have the same right hand side, a desired property to perform Barten's non-nested test. Testing models (7) and (8) assumes that there is significant structural change while testing models (2) and (5) assumes the absence of structural change. For data series where both functional form and structural change are issues, such separate tests may give different results if both functional and structural misspecification apply.

4. Development of the Compound Model and Non-nested Tests

The Rotterdam and almost ideal models are not nested. These demand systems have same right hand side, but a different left hand side. Barten (1993) suggests a compound model to nest them artificially. That is, following Barten:

$$y_{At} = x_t \beta_A + u_{it} \tag{9}$$

$$y_{Rt} = x_t \beta_R + e_{it} \tag{10}$$

where y_{At} and y_{Rt} represent, respectively, the left hand side vector of the almost ideal and Rotterdam models; x_t is a $n \times k$ matrix of exogenous variables; β_A and β_R are vectors of coefficients, specific for each system, and u_{it} and e_{it} are disturbance terms. The artificial compound model is:

$$v_{it} = \lambda e_{it} + (1-\lambda) u_{it} \tag{11}$$

or

$$v_{it} = \lambda (y_{Rt} - x_t \beta_R) + (1-\lambda)(y_{At} - x_t \beta_A) \tag{12}$$

with λ=1 representing the Rotterdam model and λ=0 representing the almost ideal model. Rewriting (12) gives:

$$y_{At} = x_t (\lambda \beta_R + (1-\lambda) \beta_A) + \lambda (y_{At} - y_{Rt}) + v_{it} \tag{13}$$

Barten argues that the difference terms y_{At}- y_{Rt} can be treated as a vector of exogenous variables and concludes that the compound model is well-specified. The null hypothesis is H_0: λ=0. If H_0 is rejected, the almost ideal model is rejected as a satisfactory explanation of reality since the alternative Rotterdam model makes a significant contribution relative to the almost ideal model. If H_0 cannot be rejected, the almost ideal model cannot be rejected by the Rotterdam model. Revising (13) gives the test for the Rotterdam model:

$$y_{Rt} = x_t(\lambda\beta_A + (1-\lambda)\beta_R) + \lambda(y_{Rt} - y_{At}) + v_{it} \tag{14}$$

Similarly, if $H_0: \lambda=0$ is rejected, the Rotterdam model is rejected by the almost ideal model. Since a system is to be tested and the estimates of λ may differ from equation to equation, the results may not be conclusive. We use two methods to apply Barten's test. With the first, the restriction $\lambda_1 = \lambda_2 = ... = \lambda$ is imposed on the system and a t-ratio test is used to test $H_0: \lambda=0$. Second, likelihood ratio (LR) tests are applied in testing the almost ideal model, viewed as the restricted model, (7) relative to the unrestricted compound model (13); and the Rotterdam, restricted, model (8) relative to unrestricted model (14). The null hypothesis is $H_0: \lambda_1=0$ and $\lambda_2=0$ in each case. For consistency with demand theory, homogeneity and symmetry conditions are explicitly imposed.

The data used in this study are quarterly retail-level prices and retail-weight aggregate Canadian consumption series for beef, pork and chicken, expressed in per capita terms, from 1967-I to 1992-III. [2] Seasonal dummy variables are specified for the last three quarters of each year. The models are estimated using the iterative seemingly unrelated regression procedure of SHAZAM 7.0.

The results of the first set of tests, designed to test only the functional form of the two systems, are given in Table 1. Viewing the Rotterdam as the alternative model, both t-ratio and LR tests indicate that the almost ideal model cannot be rejected by the Rotterdam model at the 5% significance level. The estimated

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[2] These series were made available to us by Dr. James E. Eales, Purdue University. They are derived from quarterly carcass weight disappearance series supplied by Agriculture Canada, which are adjusted to retail weights using conversion factors reported by Hewston (1987) and Hewston and Rosien (1989). The price data are derived from consumer price index measures for meats collected and published by Statistics Canada. These are converted to retail prices based on the 1986 city average retail prices for various meat cuts, aggregated by Statistics Canada's CPI weights. Turkey and fish are not included in this analysis due to their relatively minor contribution to aggregate consumption levels and consumer expenditures. These foods are not routinely consumed by all Canadian households and reliable quarterly consumption series are not available for them.

coefficient of the difference term for the beef equation is λ_1=0.35 while for the pork equation, λ_2=0.51. In the case where this coefficient is restricted to be equal within the system, the estimate is λ=0.41. None of these estimates of λ are significant at the 5% level. The likelihood ratio test of LR=3.51 is insignificant statistically (for 2 degrees of freedom, the 5% critical value of the χ^2 distribution is 5.99). Thus the almost ideal model cannot be rejected by the Rotterdam model for this data set. Viewing the almost ideal system as the alternative model, the results are inconclusive. Both the estimated coefficient of the difference term for pork λ_2 = 0.64 and the restricted λ = 0.58 are significant at the 5% level. However, for the beef equation, λ_1 = 0.49 is not significant at the 5% level. Moreover, the LR test reveals that the Rotterdam model cannot be rejected by the almost ideal model because LR=5.07 is less than the critical value of 5.99. In sum, the results for the test of functional form in the absence of structural change are inconclusive.

Table 1: Test Results for Models Without Structural Change

Alternative Model	LAIDS				ROT			
Null Model	λ_1	λ_2	λ	LR	λ_1	λ_2	λ	LR
LAIDS	--	--	--	--	0.35 (0.30)	0.51 (0.30)	0.41 (0.30)	3.51
ROT	0.49 (0.31)	0.64* (0.30)	0.58* (0.29)	5.07	--	--	--	--

Notes: λ_1 is the estimated coefficient of the difference term from the beef equation; λ_2 is from the pork equation; λ is from the system with the restriction λ_1=λ_2; LR is the likelihood ratio for which the restricted model is the linearized almost ideal model denoted by LAIDS (the Rotterdam model is indicated by ROT) and the unrestricted model is the compound model.
The figures in parentheses are standard errors; * denotes significance at the 0.05 level, based on the asymptotic t-ratio.

Before proceeding to test jointly both structural change and the two functional forms, LR tests are used to test separately the significance of incorporation of structural change for the two functional forms. Using Moschini and Meilke’s gradual switching almost ideal demand model and procedure on data for Canadian meat consumption from 1968 to 1987, Reynolds and Goddard (1991) found that t_1 =1975-I and t_2 = 1984-I gave the maximizing likelihood function. We use the findings of Reynolds and Goddard as starting points to apply the method of Moschini and Mielke to Canadian meat consumption data from the period 1967 to 1992 for both equations (7) and (8) and find that t_1=1975-I and t_2=1984-I give the maximizing likelihood function for both models. We incorporate this result in both the gradual switching almost ideal and Rotterdam models and take these models as the unrestricted models; the restricted models exclude the structural change provision. For the almost ideal model, LR=66.60 and for the Rotterdam model, LR=63.46 . For 11 degrees of freedom, the 5% and 1% critical values of the χ^2 distribution are 19.68 and 24.7, respectively. Thus, the hypothesis of no structural change in both models is rejected at the 1% significance level.

The results of the functional specification tests for the gradual switching almost ideal and Rotterdam models are reported in Table 2. Unlike the tests for the models without structural change, the results of this set of tests are conclusive. Both t-ratio and LR tests indicate that the gradual switching almost ideal model cannot be rejected by the gradual switching Rotterdam model. The gradual switching almost ideal model strongly rejects the gradual switching Rotterdam model at the 5% significance level. Moreover, the test values for λ are about 1. Barten points out that this strongly favours the alternative model. The values of the t-ratios and LR results are consistent. Comparison of the results in Tables 1 and 2 suggest that structural misspecification may affect the choice of functional form.

Table 2: Test Results for Models With Structural Change

Alternative Model	Gradual Switching LAIDS				Gradual Switching ROT			
Null Model	λ_1	λ_2	λ	LR	λ_1	λ_2	λ	LR
Gradual Switching LAIDS	--	--	--	--	-0.08 (0.56)	-0.21 (0.56)	-0.14 (0.56)	1.04
Gradual Switching ROT	1.08* (0.55)	1.21* (0.56)	1.14* (0.56)	6.27*	--	--	--	--

Notes: λ_1 is the estimated coefficient of the difference term from the beef equation; λ_2 is from the pork equation; λ is from the system with the restriction λ_1=λ_2; LR is the likelihood ratio for which the restricted model is the almost ideal denoted by LAIDS (the Rotterdam is denoted by ROT) and the unrestricted model is the compound model.
The figures in parentheses are the standard errors; * denotes significance at the 0.05 level, based on the asymptotic t-ratio.

5. The Impact of Model Choice on Elasticity Estimates

The results of the joint functional and structural specification tests suggest that the gradual switching almost ideal demand model is appropriate for analysis of Canadian meat consumption data. Statistical testing provides one criterion for model selection. The influence of model choice on elasticity estimates is also of interest. Parameter estimates for the gradual switching models, conditional on the calculated values for the switching periods prior to t_1 and following t_2, are given in Table 3 with summary statistics. The R^2 values indicate the satisfactory fit of the models. Not surprisingly the results for the gradual switching almost ideal model are close to those reported by Reynolds and Goddard (1991). The significance of the estimated parameters is similar. Both studies find a negative coefficient on the non-time varying real expenditure variable and a positive coefficient on the time-varying real expenditure term for the pork equation, suggesting that the apparent structural change has led to a large increase in the expenditure elasticity of pork.

Marshallian demand elasticities that reflect the effects of structural change, conditional on meat expenditures in the gradual switching almost ideal and Rotterdam models, are computed and reported in Table 4. The most frequently used and preferred formulae for the linearized almost ideal system (Buse, 1994) are used for this purpose. For the gradual switching almost ideal model (after structural change) these are calculated as:

$$\epsilon_{ii} = (\delta_{ii} + \pi_{ii})/w_i^a - (\rho_i + \pi_i) - 1 \tag{15}$$

$$\epsilon_{ij} = (\delta_{ij} + \pi_{ij})/w_i^a - (\rho_i + \pi_i)(w_j^a / w_i^a),\ i \neq j \tag{16}$$

$$\epsilon_{ix} = (\rho_i + \pi_i)/w_i^a + 1 \tag{17}$$

For the gradual switching Rotterdam model (after structural change) the elasticity estimates are calculated as:

$$\epsilon_{ii} = (\gamma_{ii} + \phi_{ii})/w_i^a - (\beta_i + \phi_i) w_j^a \tag{18}$$

$$\epsilon_{ij} = (\gamma_{ij} + \phi_{ij})/w_i^a - (\beta_i + \phi_i) w_j^a / w_i^a,\ i \neq j \tag{19}$$

$$\epsilon_{ix} = (\beta_i + \phi_i)/w_i^a \tag{20}$$

Table 3: Parameter Estimates for the Gradual Switching Almost Ideal and Rotterdam Models of Canadian Meat Demand

	Almost Ideal		Rotterdam	
Variable Descriptions	Beef	Pork	Beef	Pork
Non-time varying parameters for:				
Beef price	0.0742 (0.030)	-0.0118 (0.024)	-0.1721 (0.031)	0.1692 (0.024)
Pork price	-0.0118 (0.024)	0.0446 (0.024)	0.1692 (0.024)	-0.1716 (0.024)
Chicken price	-0.0624 (0.016)	-0.0328 (0.014)	0.0029 (0.016)	0.0024 (0.014)
Expenditure	0.0981 (0.053)	-0.0339 (0.046)	0.6695 (0.053)	0.2869 (0.047)
Quarter 2	0.0060 (0.003)	-0.0147 (0.002)	0.0060 (0.003)	-0.0145 (0.002)
Quarter 3	0.0117 (0.003)	-0.0215 (0.003)	0.0115 (0.003)	-0.0211 (0.003)
Quarter 4	0.0033 (0.003)	-0.0034 (0.002)	0.0029 (0.003)	-0.0029 (0.003)
Time-varying parameters for:				
Constant term	-0.0782 (0.037)	-0.3864 (0.330)	0.1740 (0.376)	-0.2191 (0.331)
Beef price	0.0784 (0.062)	-0.0679 (0.048)	0.0832 (0.062)	-0.1026 (0.048)
Pork price	-0.0680 (0.048)	0.0568 (0.046)	-0.1026 (0.048)	0.0671 (0.047)
Chicken price	-0.0104 (0.031)	0.0112 (0.024)	0.0194 (0.032)	0.0356 (0.028)
Expenditure	0.0061 (0.081)	0.0806 (0.071)	-0.0495 (0.081)	0.0432 (0.071)
Quarter 2	0.0161 (0.004)	-0.0153 (0.004)	0.0155 (0.004)	-0.0158 (0.003)
Quarter 3	0.0103 (0.005)	-0.0051 (0.004)	0.0102 (0.005)	-0.0053 (0.004)
Quarter 4	0.0002 (0.004)	0.0091 (0.004)	0.0008 (0.004)	0.0085 (0.003)
R^2	0.72	0.83	0.92	0.82
D-W	2.45	2.45	2.44	2.44
ML value	755.8		753.2	
n	103		103	

Note: Figures in parentheses are standard errors.

Table 4: Marshallian Elasticities of Demand Calculated from Almost Ideal and Rotterdam Models

	Price of			Expenditure
Elasticities of:	Beef	Pork	Chicken	
Almost Ideal Model				
Before Structural Change				
Beef	-0.965	-0.079	-0.131	1.176
	(0.068)	(0.058)	(0.037)	(0.095)
Pork	0.021	-0.832	-0.087	0.898
	(0.096)	(0.093)	(0.045)	(0.141)
Chicken	-0.241	-0.103	-0.0725	0.417
	(0.173)	(0.153)	(0.162)	(0.239)
After Structural Change				
Beef	-0.797	-0.221	-0.191	1.209
	(0.105)	(0.078)	(0.056)	(0.107)
Pork	-0.358	-0.694	-0.110	1.161
	(0.151)	(0.125)	(0.076)	(0.163)
Chicken	0.009	0.101	-0.412	0.301
	(0.115)	(0.084)	(0.110)	(0.127)
Rotterdam Model				
Before Structural Change				
Beef	-0.978	-0.095	-0.127	1.201
	(0.068)	(0.059)	(0.032)	(0.096)
Pork	0.028	-0.803	-0.088	0.863
	(0.097)	(0.093)	(0.046)	(0.142)
Chicken	-0.194	-0.109	-0.091	0.395
	(0.177)	(0.156)	(0.166)	(0.244)
After Structural Change				
Beef	-0.799	-0.225	-0.225	1.248
	(0.106)	(0.078)	(0.057)	(0.107)
Pork	-0.338	-0.649	-0.158	1.148
	(0.151)	(0.125)	(0.074)	(0.164)
Chicken	-0.012	0.109	-0.329	0.231
	(0.117)	(0.085)	(0.113)	(0.129)

Note: Figures in parentheses are standard errors.

Elasticity estimates that reflect demand responses before structural change are obtained from (15) to (20) by setting π=0 in the gradual switching almost ideal model, setting ϕ=0 in the gradual switching Rotterdam model, and substituting w_i^b for w_i^a where superscripts a and b denote, respectively, the period after and before structural change. Thus w_i^a is calculated as the mean budget share for good i over the period before structural change and w_i^b is the mean budget share of i after structural change.

After structural change, the expenditure elasticity increased for pork and beef and fell for chicken. The absolute value of the own price elasticity of demand increased for chicken (from - 0.07 to -0.41) and fell for pork (from -0.83 to -0.69). The relationships for the pairs of meats, chicken and beef, and chicken and pork changed from complementarity to substitution. It is clear that the structural change specification has a considerable effect on estimated elasticities. The bias of structural change, which reflects the effects of structural change on quantity demanded with prices and expenditure levels held constant, can be calculated at the sample means of the exogenous variables, following Moschini and Meilke (1983). These calculations, made for the gradual switching almost ideal model, are reported in Table 5. As Reynolds and Goddard (1991) found, these results show that structural change in Canadian meat consumption has been significantly biased toward chicken. The results of Table 5 suggest a 46 percent budget share increase for chicken could have been expected, had prices and total expenditures remained constant following structural change. However, in contrast to pork prices, chicken prices increased continuously and the actual shift in mean budget shares for chicken was only 10.6 percent. We find negative but insignificant structural change bias for pork and beef, whereas a significant bias against beef and neutral effect for pork were reported by Reynolds and Goddard for the shorter time period to 1987. This is consistent with the considerable variability over time in Canadian pork consumption that is associated with cyclic changes in pork production and prices; no significant trend in pork prices has been evident since the mid-1980s whereas beef prices tended to increase.

Table 5: Bias of Structural Change

Meat	Estimated Bias	Actual Mean Expenditure Shares		Actual Share-Shifts
		Before Structural Change $(\bar{w}^b)$	After Structural Change $(\bar{w}^a)$	$(\bar{w}^a - \bar{w}^b)$
Beef	-0.0716 (0.377)	0.557	0.496	-0.061
Pork	-0.3893 (0.331)	0.333	0.287	-0.046
Chicken	0.4609 (0.189)	0.110	0.216	0.106

Note: Figures in parentheses are standard errors. The estimated structural change bias is calculated at the sample means as: $B_i = \tau_i + 0.25(\Sigma_k \tau_{ik})$.

In contrast to the structural change specification, the effect of the choice of the two functional forms on demand elasticities is relatively small. As shown in Table 4, the two models generate very similar numeric estimates of elasticities despite the differences in their behavioural implications and in their consistency with the data. Most previous studies of the demand for meat that have used the almost ideal and Rotterdam models have also found only minor differences in the elasticity estimates generated by the two models. For example, Wohlgenant and Hahn (1982) used both models and found only minor differences in elasticities for US meats. This was also the conclusion of Alston and Chalfant (1993). Even so, the apparently small effect of functional form choice on elasticities of demand does not mean that choice of an appropriate system is unimportant, since the effect on elasticities is not known *ex ante* and the purpose of demand analysis may not be limited to elasticity calculation.

6. Conclusions

Two sets of non-nested tests are applied in this paper: one is for two commonly used models without structural change imposed and the other is for these models with structural change imposed. For Canadian meats, the non-nested test for the almost ideal and Rotterdam models without structural change is inconclusive and neither model can be rejected statistically. However, the non-nested test for the two functional models with structural change is conclusive.

We find that the alternative functional specifications of the linear almost ideal and Rotterdam models do not greatly affect estimated elasticities for data on Canadian meat consumption. These models produce very similar elasticities although the latter model is rejected when structural change is incorporated. The finding that functional specification has a minor effect on elasticities was also the conclusion, based on data for US meat consumption, by Wohlgenant and Hahn (1982) and Alston and Chalfant (1993). However, Lee et al. (1994), concluded that the Rotterdam model gave questionable elasticity estimates for data on general consumption patterns in Taiwan. The general relationship between functional choice and demand elasticities for the almost ideal and Rotterdam models evidently may vary with the data set. In contrast to the influence of the two functional forms on the elasticities of demand for meats, the effect of structural change on the estimates of demand elasticities for meats is appreciable. When both issues are considered jointly, the almost ideal model with structural change performs best for data on Canadian meat consumption. Specifically, we find that the gradual switching Rotterdam model is rejected by the gradual switching almost ideal model at the 5% significance level, indicating that the almost ideal model with a gradual transition structural change function better explains Canadian meat demand than does the gradual switching Rotterdam model.

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