ACCOUNTING FOR CHANGES IN DEMAND

Julian M. Alston and James A. Chalfant

Abstract

Parametric tests for structural change are conditional on the joint hypothesis of functional form and other aspects of the model specification. This problem is often disregarded in demand analysis. Using quarterly Australian per capita consumption data on meat (beef, lamb, pork, and poultry) for 1970:2 to 1988:4, we found consistency with stable preferences using nonparametric tests, highly significant seasonality and trends with the Lewbel, AIDS and translog models, and significant seasonality but no trends with a Rotterdam model. The specification differences did not affect elasticities much. The trends are quite large and economically important.

Monte Carlo evidence indicates that functional form specification errors can lead to substantial increases in the probability of finding trends when they are not present in the data generating mechanism; but specification error also reduces the probability of finding trends that are in the data. The AIDS model had a high probability of finding trends whether they were present or not; the Rotterdam model had a lower probability of finding trends whether they were present or not. Previous studies have questioned the power of nonparametric tests. The power and significance of parametric tests for structural change is also called into question.

Key Words: functional forms; regression diagnostics; structural change; meat demand.

Invited paper presented to the 35th Annual Conference of the Australian Agricultural Economics Society, Armidale, February 11-14, 1991

Julian M. Alston is an Associate Professor in the Department of Agricultural Economics, University of California, Davis, CA 95616, USA and James A. Chalfant is an Associate Professor in the Department of Agricultural and Resource Economics, University of California, Berkeley, CA 94720, USA.

ACCOUNTING FOR CHANGES IN DEMAND

And [is] every failure of an economic prediction a showing that tastes have changed? Not quite. Changing tastes can invalidate a prediction, but so too can a mistaken analysis -- the neglect of some other important variable, the improper formulation of the theory, clumsy statistical technique. The economist looks for these sources of failure much more often than he looks for changing tatses -- simply because they have been found to be more probable than changing tastes.

-- G.J. Stigler (1966, p.39)

Further, the specification error leaves no trace of its evil presence. The methods of statistics deal, quite properly, with sampling errors alone, and cannot be expected to help with others. When these others are serious, as we all believe they usually are, there is no point in continuing to talk relatively much about the (relatively small) sampling error. To do so, as economists and other quantifiers do nowadays on a massive scale, is to imitate the drunk who looks for his wallet under the lamp post because the light is better there.

-- D.N. McCloskey (1985, p.162)

It appears to be widely believed that there has been a significant change in consumer preferences away from red meat during the 1970s and 1980s. Certainly, per capita consumption of beef has declined and per capita consumption of chicken has risen. The most common explanation in the literature (and in discussions we have had) is that consumers have become more aware of and concerned about the health consequences of diet and they believe that "white meat" (especially chicken and fish) is healthier. The industry has responded by investing in new product development and promotion campaigns in an attempt to restore demand for red meat. This is not just an Australian phenomenon. Similar changes in consumption patterns have occurred, for instance, in the United States and Canada, and the industry has responded similarly in those countries. Agricultural economists have not neglected this question. There have been a large number of studies that have attempted to test for and measure structural change in meat demand.[1] The results of those studies are mixed and opinion remains divided.

Our interest in this area began with a paper presented at the Australian Agricultural Economics Society annual conference five years ago (Chalfant and Alston, 1986). Our contention then and in a series of subsequent papers has been that the typical structural change test is a joint test of the hypothesis of stable preferences and all other aspects of the specification; therefore specification errors could lead to false findings of structural change. In an article that grew from our 1986 conference paper (Chalfant and Alston, 1988) we used a *nonparametric* testing approach that avoided the problem of specification as a joint hypothesis. Using a test developed by Varian (1982, 1983), we found that U.S. and Australian per capita meat consumption data could have been generated from stable preferences and that, therefore, any conclusions from those data that tastes have changed must derive from restrictions on the nature of the demand equations (e.g., to be of the almost ideal form).[2]

One important *caveat* in that study was the unknown "power" of the nonparametric test: Is the test capable of detecting changes in tastes in significantly trended data?[3] We suggested imposing prior beliefs about elasticities as a way of increasing power (as opposed to imposing an *ad hoc* functional form). In a more recent paper (Alston and Chalfant, 1991a) we addressed the power question more systematically. Using Monte Carlo experiments we found that the probability of detecting a structural change with that test was rather low (say 25 to 50 percent for a change of the type and magnitude suggested by Moschini and Meilke (1989) for U.S. meat consumption data). In that study we also showed that typical *parametric* tests for structural change were highly susceptible to specification error. Specifically, errors in specification of the functional form for demand equations led to significant evidence of autocorrelation and discrete demand shifts (Chow tests) when they were not present in the data generating mechanism. In addition, while Canadian meat consumption data were consistent with stable preferences according to both the nonparametric test and the Rotterdam model, using an almost ideal demand system (AIDS) led to significant time trends, Chow tests, and autocorrelation.

Those results led us to undertake to some further work on Chow tests. In a third paper (Alston and Chalfant, 1991b) we used Monte Carlo methods to explore in more detail the implications of functional form specification error for structural change tests. We found that apparently inoccuous specification errors could lead to quite dramatic increases in the probability of finding significant autocorrelation or a significant Chow test. Further, the "maximum" Chow test was found to be quite unreliable. These results held in a range of applications including single equation models of investment behavior and meat consumption and integrable demand systems applied to artifical data.

To summarize, the typical parametric structural change tests (which in other contexts are used as broader diagnostic tests) seem sensitive to choices about the functional forms for demand equations. Parametric models using Australia's meat consumption data (e.g., Beggs, 1988;

Martin and Porter, 1985; Chalfant and Alston, 1986) provide some (albeit weak) support for the structural change hypothesis. On the other hand, nonparametric tests with the same data have not supported the structural change hypothesis but, given the Monte Carlo evidence on power of that test, it is not clear how much weight should be attached to that result. If we believe the nonparametric result, a question that remains is what is the nature of the stable system of well-behaved meat demand equations?

As well as having implications for structural change tests, specification error can have implications for elasticity estimates and other questions, such as whether demand is characterized by dynamics or seasonality. The purpose of this paper is to illustrate these ideas using Australia's meat consumption data. The emphasis will be on the functional form for demand equations, though presumably other specification choices will have similar implications. We will explore the effects of functional form on various results of interest in demand studies.

The Demand Models

There have been many attempts to estimate demand relationships for meat in Australia but to our knowledge only one of these studies (Murray, 1984) has successfully estimated the demand for meats in a systems approach using the restrictions from consumer preference theory.[4] Murray estimated a total of ten alternative specifications and rejected all but three of those using nested hypothesis tests. Murray also rejected the remaining three models -- the almost ideal demand system (AIDS), the translog and the indirect addilog -- using non-nested hypothesis tests, but she noted the unknown properties of the Cox test in small sample nonlinear systems. In conclusion, Murray suggested the use of alternative data, such as quarterly time series, and the incorporation of dynamic components into the models, to determine more conclusively the validity of static utility theory in this area of research.

In this study we treat meat as a weakly separable group in which consumption of individual meat items depends only on expenditure on the group and the prices of goods within the group. The meat group is assumed to comprise four commodities: beef, lamb, pork, and poultry. We study the demand for this group of commodities using quarterly data for the period 1970:2 to 1988:4. These data are similar to those of Martin and Porter (1985) but they cover a more recent time period and mutton is excluded.[5]

As a first step, we applied the nonparametric test to these new data. There were no violations of the Generalized Axiom of revealed Preference (GARP), so the data for the four meats excluding mutton are consistent with having been generated by a stable system of well-behaved per capita demand equations.[6] A stable system of demand equations will not exhibit any trends or, in the context of quarterly data, seasonal patterns. Thus, this nonparametric result contradicts the common finding that quarterly meat demand equations are characterized by trends and seasonality.[7] The question that this raises is a familiar one: Is the nonparametric result a reflection of low power of the tests or is the parametric result due to specification error?

In order to pursue this question, we consider four alternative demand models. Two of these, the almost ideal demand system (AIDS) and translog, were among Murray's preferred models.[8] Our third model is Lewbel's (1989) model that nests the AIDS and translog models as special cases. The fourth is the absolute price version of the Rotterdam model.[9] These four demand systems are represented in equations (1) through (4).

Translog Model:

$$S_i = \frac{\alpha_i + \sum_{j=1}^{n} \beta_{ij} \ln\left(\frac{P_j}{M}\right)}{\sum_{k=1}^{n} \alpha_k + \sum_{k=1}^{n}\sum_{j=1}^{n} \beta_{kj} \ln\left(\frac{P_j}{M}\right)} \tag{1}$$

Almost Ideal Demand System (AIDS):

$$S_i = \alpha_i + \sum_{j=1}^{n} \gamma_{ij} \ln P_j + \beta_i \ln\left(\frac{M}{P}\right) \tag{2}$$

where $\ln P = \alpha_0 + \sum_{k=1}^{n} \alpha_k \ln P_k + \frac{1}{2} \sum_{k=1}^{n} \sum_{j=1}^{n} \gamma_{kj} \ln P_k \ln P_j$.

Lewbel's Nested AIDS and Translog Model:

$$S_i = \frac{\alpha_i + \sum_{j=1}^{n}\gamma_{ij}\ln P_j + \beta_i \ln\left(\frac{M}{P}\right) - \left(\sum_{k=1}^{n}\gamma_{ki} - \beta_i \sum_{j=1}^{n}\sum_{k=1}^{n}\gamma_{jk}\ln P_k\right)\ln M}{1 + \sum_{j=1}^{n}\sum_{k=1}^{n}\gamma_{jk}\ln P_k} \tag{3}$$

$$\text{where } \ln P = \alpha_0 + \sum_{k=1}^{n}\alpha_k \ln P_k + \frac{1}{2}\sum_{k=1}^{n}\sum_{j=1}^{n}\gamma_{kj}\ln P_k \ln P_j.$$

Rotterdam Model (Absolute Price Version):

$$\bar{S}_i \Delta\ln q_i = \sum_{j=1}^{n}\gamma_{ij}\Delta\ln P_j + \beta_i \Delta\ln\hat{Q} \tag{4}$$

$$\text{where } \Delta\ln\hat{Q} = \sum_{k=1}^{n}\bar{S}_k \Delta\ln q_k \text{ and } \bar{S}_{i,t} = \frac{1}{2}(S_{i,t} + S_{i,t-1}).$$

In these equations, q_i = per capita consumption of meat type i, P_i = per unit price of meat type i, M = total expenditures per capita on the four types of meat (i.e., $M = \Sigma_i P_i q_i$), and S_i = the share of meat type i in per capita expenditure on meat (i.e., $S_i = P_i q_i / M$). Using these four functional forms we estimate Australia's meat demand equations and test for time trends and seasonal shifts in demand and we evaluate their economic importance. For each plausible model we compute demand elasticities and compare them among models.

The four models are "locally flexible" functional forms in that there are no restrictions implied by the function upon the values that elasticities may take at any data point. The flexibility of these models to allow elasticities to vary with data is limited only when the parameters are fixed. The Rotterdam model is not nested with any of the other models.[10] The AIDS and translog models are special cases in the Lewbel model and can be imposed and tested as the following parametric restrictions (as can be seen in equation (3)):

Translog: $\beta_i = 0 \;\forall\, i.$ AIDS: $\sum_{k=1}^{n}\gamma_{ki} = 0.$

Because expenditure shares sum to one, the covariance matrix for the full system of equations is singular and one equation is deleted for estimation. When we estimate by iterated nonlinear seemingly unrelated regressions methods, we obtain maximum likelihood estimates of the parameters and the results are invariant to the choice of which equation to delete (Barten, 1971). Throughout, we delete the equation for poultry in the estimation and we use the adding up restrictions to recover its parameters. In all of our estimations we impose homogeneity, adding up and symmetry restrictions as maintained hypotheses.[11]

Some Estimation Issues

Two issues that arose in the estimation warrant some discussion. First, what is the role of the intercept in the AIDS price index (α_0 in equations (2) and (3)), how should that parameter be estimated, and what problems does it create? Second, what is the appropriate way to introduce time trends and seasonality into share equations and what are the implications of those choices?

The AIDS has been very popular in recent studies, but in most cases the "linear approximate almost ideal demand system" (LA/AIDS) is used as an approximation. The LA/AIDS model uses Stone's price index (P*) instead of the AIDS index (P in equation (2)).

Stone's Price Index is: $\ln P* = \sum_{i=1}^{n} S_i \ln P_i$.

The same type of approximation is also possible with Lewbel's more general model. Using Stone's price index instead of the AIDS price index in equation (3) yields the approximate Lewbel model, which nests the translog and the linear approximate AIDS model. We use the "true" AIDS model for most of the work below. Mainly for illustrative purposes, we also try the approximate forms.

Deaton and Muellbauer (1980a) pointed out that estimating α_0 is likely to be troublesome. One way they suggested to avoid this problem is to use Stone's index as an approximation to the AIDS price index. The drawback of this is that the linear approximate model is not an integrable demand system. Another way is to fix the parameter prior to estimation. Deaton and Muellbauer (1980a, p. 316) suggested that "Since this coefficient can be interpreted as the outlay required for a minimal standard of living when prices are unity (usually in the base year; see the Appendix), choosing a plausible value is not too difficult." What this seems to overlook is that rescaling the price data, to create indices that equal one in a base year, affects the values of the expenditure shares and thereby affects the model and its parameters in a substantive way. The third alternative is to estimate the full model with α_0 as a free parameter.

Our conclusion from trying all three of these approaches in both the Lewbel model and the AIDS model is that neither of the first two solutions is a good one. The third option did, in fact, prove to be somewhat troublesome: we learned that estimating α_0 may indeed be difficult and leaving it free meant that nonlinear estimation could be slow and the results could be sensitive to starting values. However, we did succeed in estimating the parameter and, after much experimentation with starting values, we believe we found a global maximum in every case. The lesson was not to abandon the estimation of α_0, but, rather, to be conscious of the hazards and do much sensitivity analysis.

In relation to incorporating trends, previous studies using similar data have indicated shifts over time in per capita demands for meat using models of the types being tried here with both Australian data (e.g., Alston and Chalfant, 1987; Chalfant and Alston, 1986) and U.S. data (e.g., Moschini and Meilke, 1989 and Moschini, 1991). Testing for significance of trends in demand equations is one approach we use to test for structural change (or to test specifications given a maintained hypothesis of stable demand).

We considered two alternative ways to introduce the effects of time (trends and quarterly dummy variables) into the models. One is to incorporate shift variables as modifications of the intercept terms already in the models (α_i's) and to preserve the adding up restrictions on those modified intercepts. This method ensures that the augmented model is compatible with theory but it means that the effects of time are confounded with the other components of the model in some way that is not straightforward to disentangle. The alternative is simply to add terms linearly to the model rather than introduce them as modifications of the intercepts already in the models. This approach will cause the model to violate symmetry restrictions, but it is more straightforward to interpret. We opted for the former method. In each of the first three equations this amounts to modifying the intercepts according to:

$$\alpha_i = \alpha_{0i} + \tau_i T + \sum_{k=1}^{n} \theta_{ik} QD_k$$

where T is a time trend set equal to 1 in 1970:2 and QD_k (k=1,2 or 3) are quarterly intercept dummies. The full effect of trends and seasonality on shares is complicated since the intercept terms (the α_i's) also appear in the AIDS price index.

In demand models that have quantities (or logarithms of quantities) as dependent variables, the interpretation of seasonality and time trends is straightforward and intutive. This is not so when share equations are estimated with quarterly intercept dummies and trends, especially when the intercepts enter in a nonlinear fashion. In the Rotterdam model, the interpretation is perhaps more difficult. In the Rotterdam model we can include an intercept in the equation for each good, and doing so implies changes in consumption over time (e.g. Theil and Clements, 1987), but it is not a straightforward trend -- it is a trend in the term $S_i \Delta \ln q_i$. Similarly, quarterly intercept dummy variables will capture seasonal changes in consumption but it is seasonality in the term $S_i \Delta \ln q_i$. To represent seasonality in the Rotterdam model, we include four quarterly intercept dummy variables in each equation with a restriction that the co-efficients sum to zero within the equation. Thus the augmented Rotterdam model is:

$$\bar{S}_i \Delta \ln q_i = \tau_i + \sum_{k=1}^{4} \theta_{ik} QD_k + \sum_{j=1}^{n} \gamma_{ij} \Delta \ln P_j + \beta_i \Delta \ln \hat{Q}, \quad with \sum_{k=1}^{4} \theta_{ik} = 0. \tag{5}$$

With these specifications, the test for seasonality is a test of whether the nine co-efficients on the quarterly dummies (θ_{ik}, where i,j = 1, 2, or 3) are jointly significant, the test for significant trends is a test of whether the three coefficients on trend (τ_i, where i = 1,2, or 3) are jointly significant and the test for seasonality and trends is the joint test of those two hypotheses.

Coefficient Estimates and Hypothesis Tests

The 16 alternative models (i.e., Lewbel, translog, AIDS, and Rotterdam, with and without the quarterly dummies and trends) were estimated using 75 quarterly observations (1970:2 to 1988:4). The coefficient estimates for the four models, in the case where trends and quarterly intercepts were included, are reported in Table 1. Also included are the parameters from the Rotterdam model without trends; it is notable that excluding the trend variables had little impact on the other coefficients. Blanks in that table indicat coefficients that were not estimated directly but were calculated from restrictions.

In order to test for statistical significance of trends and seasonality, the same models were estimated without trends, without seasonality, and without both trends and seasonality. The Rotterdam model and the translog model were easy to estimate in every case, but the Lewbel and AIDS models were relatively difficult to estimate and we attribute these difficulties to the α_0 parameter (when we fix that parameter the models are much easier to estimate). The estimations seemed otherwise satisfactory in every case. We repeated the same estimations using the approximate version of Lewbel's model.

Tables 2, 3 and 4 summarize the results from hypothesis tests. In Table 2, tests for statistical significance of trends and seasonality indicate that seasonality is highly significant in every model and that trends are highly significant in every model except for the Rotterdam model. This is consistent with some previous results with Canadian data (Alston and Chalfant, 1991a). In the Rotterdam model, trends were not statistically significant regardless of whether seasonal dummies were included. The results using the linear approximation to the AIDS price index, in Table 3, are similar: trends and seasonality are highly significant in these models as well. These results indicate that, for the Lewbel type models (including the AIDS and translog and the LA/AIDS), trends and seasonality should be included.

Table 4 reports the results of tests for AIDS and translog restrictions on the Lewbel models with and without trends and seasonality. The translog restriction is rejected in every case except for the case when both trends and seasonality are excluded, and the results in Table 2 indicate that that model is misspecified. In contrast, we rejected the AIDS restriction in only one case, the model with trends but excluding seasonality. However, when the linear approximation is used (i.e., Stone's price index instead of the AIDS price index), the LA/AIDS model is comprehensively rejected in every case by the LA/Lewbel model. The translog model is rejected in every case except for the case when seasonality is excluded and the results in Table 3 indicate that that model is misspecified. While we have not tested the linear approximation formally, we are inclined to prefer the model based on the true AIDS price index.

Expenditure elasticities and uncompensated price elasticities are reported in Tables 5, 6, and 7. The numbers in these tables are the means of elasticities that were computed at every data point. Table 5 shows elasticities from the Lewbel, translog, AIDS and Rotterdam models when trends and seasonality were included. The elasticities are remarkably similar among the four models and, on the whole, plausible. It is notable that the expenditure elasticity for poultry is very small in every model (negative in three of the four models) and that the own-price elasticities from the Rotterdam model tended to be smaller than those from the other models, especially for poultry. Table 6 shows the corresponding elasticities when the trends are excluded.

In the Rotterdam model trends were statistically insignificant; in the other three models trends were highly significant. In the Rotterdam model, leaving out trends had a negligible effect on the elasticities. In the other three models, leaving out trends had some impact on the elasticities. In particular, the expenditure and own-price elasticities of demand for poultry are much larger (in absolute value) when trends are left out from those models. This should not be surprising. Previous studies have found estimating poultry demands troublesome with these data (e.g., Fisher, 1979; Martin and Porter, 1985). Similar results were obtained by Chalfant and Alston (1986): poultry elasticities were highly sensitive to inclusion of trends. We suspect that these problems are due to the high correlation of poultry price, poultry consumption and total expenditure with time.

In summary, three models are not rejected -- the Rotterdam model without trends, the Lewbel model including trends, and the AIDS including trends, all with seasonal dummy variables.[12] The elasticities computed from these three models are virtually identical. When trends are excluded the elasticities vary much more among the (misspecified) models.

Interpretation of Results

It is not easy to draw crisp conclusions from the mixture of results here. In relation to the question of changes in tastes, the results are quite mixed. The nonparametric results do not reject the hypothesis of stable preferences, nor does the Rotterdam model, whereas the Lewbel model (and its special cases) do. We know that a misspecified model can lead to false findings of structural change using Chow tests (Alston and Chalfant, 1991b); but we suspect that some specification errors might also mask structural change. We also suspect that the nonparametric test might have low power. On balance, we are inclined to conclude that we cannot reject the hypothesis of constant preferences with these results -- but we cannot be confident that demands have not changed. This is because we do not have any basis for believing that any of the models, including the Rotterdam model, is the true data generating mechanism (or a good approximation). At most these models include 16 free parameters to capture all of the effects of relative prices and total meat expenditure on per capita consumption over the 19 year period. It should be surprising if such parametrically parsimonious models were not misspecified. This could be tested by fitting more flexible models (e.g., the Fourier type models of Gallant, 1981; and Chalfant, 1987), but this process is potentially endless.

One check on the models is their elasticities (in Table 5). At least in that regard the models are in close agreement and the results are plausible. It could be that this is due to the similarity of the models and that a really different model (e.g., a generalized Leontief, or a quadratic expenditure system) would imply much different elasticities. The models that survived the tests above all imply that the four meats are net substitutes (i.e., compensated cross-price elasticities of substitution were all positive) and negativity was satisfied (i.e., all of the compensated and uncompensated own-price elasticities were negative). The one jarring note is the uncommonly low estimate of the elasticity of demand for poultry with respect to total meats expenditure which suggests that poultry may be an inferior good. On the other hand, we cannot reject that possibility based on prior beliefs. One lesson from these results is that it is worth checking some specification choices. With these data the LA/AIDS model seems to be a poor choice, especially if estimated without trends and seasonality. Indeed, any of the models (except the Rotterdam model) estimated without trends and seasonality would be misspecified and the elasticities could be misleading.

A second check on the results is to evaluate the importance of the trends found by the various models. This is not straightforward because the trend coefficients enter the equations in an awkward fashion (as intercept changes) in order to preserve the properties of the models. In the Lewbel-type models, the trends in the intercepts were statistically significant. We solved the share equations for the effects of the trend term on the quantity of each meat type and then calculated those effects at every data point.[13]

Table 7 includes the mean change in quantity per quarter attributable to trend for each of the four goods for each model. In the case of the Rotterdam model, the trend effects are not strictly comparable to the others for two reasons. First, trends were not statistically significant in the Rotterdam model and thus the best estimate we have for the trend effects is zero rather than the figures shown. Second, the effect in the Rotterdam case is calculated as the effect at the mean rather than the mean of the effects across the sample. The contrast with the other models is interesting, but not too much should be made of it.

The effects were virtually identical among the other three models. Consider the Lewbel results. For beef, the trend implied an average *reduction* per quarter in quarterly per capita consumption of 26.6 grams (0.22 percent of average quarterly consumption). Over the 75 quarters this implies a total reduction of 2 kilograms in quarterly per capita consumption (17.2 percent of average quarterly consumption). For lamb, the trend implied an average *reduction* per quarter in quarterly per capita consumption of 10.1 grams (0.24 percent of average quarterly consumption). Over the 75 quarters this implies a total reduction of 0.76 kilograms in quarterly per capita consumption (18.0 percent of average quarterly consumption). For pork, the trend implied an average *increase* per quarter in quarterly per capita consumption of 15.4 grams (0.41 percent of average quarterly consumption). Over the 75 quarters this implies a total increase of 1.15 kilograms in quarterly per capita consumption (31.0 percent of average quarterly consumption). Finally, for poultry, the trend implied an average *increase* per quarter in quarterly per capita consumption of 29.8 grams (0.67 percent of average quarterly consumption). Over the 75 quarters this implies a total increase of 2.24 kilograms in quarterly per capita consumption (50.2 percent of average quarterly consumption).

These are dramatic effects. Clearly, if we were to believe that these figures for trend effects are meaningful estimates of the effects of changes in tastes -- or, indeed, changes in per capita demand from any other source such as demographic change -- the results would be very important economically, too. Even the strongest protagonists of the taste change hypothesis have not suggested trends as important as these in previous studies. For example, with U.S. data for a similar time period, Moschini and Meilke suggested structural change accounted for a 6 percent decline in beef share, and increases of 0.3 percent in the share of pork, 8.6 percent in the share of chicken, and 11 percent in the share of fish.

Therefore, it seems prudent to take our measures of trends in consumption with a grain of salt. This is more clearly apparent when the quarterly values of trend-induced changes in consumption are plotted against time and against prices. The changes in quantities (resulting from trends in shares) are very highly correlated with prices; it is disturbing to note that the computed shifts away from beef were most pronounced when beef prices were lowest. This is a further signal that trends in share equations are difficult to translate into meaningful measures of structural change. It adds to our concerns about the appropriate form for trends to represent structural change in models of these types.

Putting aside concerns about their validity and interpretation, the results in Table 8 reinforce the point that structural change in demand could be of great economic importance and that, therefore, having plausible results that can be used with confidence would be valuable. With that in mind, we carried out some Monte Carlo experiments to get a sense of the implications of specification choices for finding trends in demand models.

Monte Carlo Results

The Monte Carlo experiments that follow are designed to examine how use of the wrong model might lead to an incorrect conclusion from a structural change test -- a test for trends in the model.[14] This requires a comparison to the correct model, so a means to generate data for Monte Carlo experiments is needed. It is easy to generate data consistent with the assumed structure of any estimated regression model. Suppose the model $y = X\beta + \epsilon$ is estimated with a sample of size T, with ϵ assumed to be a vector drawn from $N(0, \sigma^2 I_T)$. Then new vectors y^* can be created as:

$y^* = X\hat{\beta} + \hat{\sigma} e$, where

e is a random draw from the standard normal distribution,

$\hat{\beta}$ is the sample estimate of β, and

$\hat{\sigma}$ is the sample estimate of σ.

A regression of y^* on the X matrix is a correctly specified model. Thus, the usual structural change test has properties that are known from the theory -- using a chosen level of significance α, the test will falsely indicate structural change $100^*\alpha$ percent of the time. Now suppose that the wrong functional form is estimated. Instead of y* on X, the model is estimated with transformations of the data, such as logarithms or square roots. How often will the test for structural change indicate that it has occurred?

We considered a simple data-generating mechanism that has been used previously in Monte Carlo studies of demand systems by Kiefer and MacKinnon (1976) and Wales (1984). We made use of the data-generating mechanism used by Wales, who had adopted the one from Kiefer and MacKinnon with only slight modification. Briefly, the setup is as follows. The *Linear Expenditure System* (LES) is the system of share equations that arises from Stone-Geary type preferences:

$$S_i = \frac{P_i b_i}{M} + a_i \left(1 - \sum_{k=1}^{n} \frac{P_k b_k}{M}\right) + e_i .$$

The b_i parameters are interpreted as pre-committed or subsistence quantities of the goods consumed, while a_i denotes the marginal budget share of good i. Hence, budget shares are a linear function of the 'supernumerary' income, the amount of income that remains after the precommitted quantities have been purchased. As is well known, the LES satisfies all of the restrictions from consumer theory (provided that the a_i's are positive) since it is derived from well-behaved preferences.

Wales generated data using the data for prices and income given in the appendix of Kiefer and MacKinnon's paper with the parameter vectors chosen to be a = (0.2 0.4 0.4)′ and b =(0.2 0.1 0.3)′. The prices and income generated by Kiefer and MacKinnon consist of 40 observations that they considered to be consistent with the sort of data typically encountered -- trends in relative prices and income.

With data for prices and income and the assumed parameter values, it is possible to generate a series of predicted quantities for 3 goods over 40 observations; data sets with a stochastic component can be generated by appending an error term. Following Kiefer and MacKinnon and Wales, error terms were generated to exhibit timewise independence but contemporaneous correlations by generating vectors of N(0,1) random variables and transforming them to have the covariance matrix:

$$\Sigma = \begin{bmatrix} 0.000036 & -0.000025 \\ -0.000025 & 0.000049 \end{bmatrix}$$

Errors for the first two equations were generated in this fashion and errors for the third equation were generated as $e_3 = -(e_1 + e_2)$ so that the data would satisfy adding-up, etc., by construction. We generated 250 replications of linear-expenditure system data in this manner.

First we fit the correct (LES) model to these data and tested for structural change by simply adding linear trend terms to the LES share equations. Under the assumption that the parameters in Σ are unknown, but constant across the data, we tested structural change in the LES by testing whether the trends in all three equations (two independent parameters) are simultaneously zero using a Wald test. Then we performed the same test with two incorrect models, the IDS and the Rotterdam model.

Concerns over the behavior of tests for structural change and the use of an estimated, rather than a known covariance matrix have been raised in several papers.[15] For that reason, we examined the behavior of the tests when the true LES model is estimated as well as when the misspecified models are tried. This allows us to determine the extent to which a bias toward (or away from) rejection of hypotheses is due to the behavior of the test statistic itself, and the extent to which specification error has influenced its performance.

The results indicated a small amount of bias toward rejection of true hypotheses, consistent with Wales (1984), Laitinen (1978), and other previous studies, even when the correct model was estimated. In 250 replications, the Wald test was significant in 18 cases (7.2%) at the 5% level and in 8 cases (3.2%) at the 1% level. Thus, there is some justification for concern over the behavior of test statistics, and one might want to adjust the rejection region along the lines suggested by Laitinen and others.

When we estimated the LA/AIDS using the same data, the rejection probabilities increased substantially. Using a 5% critical value, the absence of trends was rejected in over one quarter of the replications (69 of 250 replications, a frequency of 27.6% rather than 5%) and the proportion of rejections fell only to 11.2% (28 of 250 replications) when the 1-percent value was used.[16] Functional form errors clearly are primarily responsible for these high frequencies of false rejections. The tendency of the Wald test to over-reject cannot account for these high frequencies of false rejections and the size corrections would not correct the problem[17].

In contrast, when the Rotterdam model was estimated, the frequency of false rejections of the stable model was much smaller than the nominal size of the tests. Using a 5% critical value, the absence of trends was rejected only once in 250 replications (a frequency of 0.4% rather than 5%) and there were no rejections at all when the 1-percent value was used. Here we have a case where functional form errors seem to be responsible for low frequencies of false rejections.

To consider further the question of power of these tests, we tried the same experiment with trends introduced into the data generating mechanism. First we introduced a simple additive trend term in the LES, scaled as $\delta = [-0.001, 0.0005]$ so that the total decline in the first share over 40 observations would be about 0.04, and the other two shares each would rise about 0.02.[18] When this was done the frequencies of correct rejections for the three models rose, with a 5% test, to 42.4% for the LES, 30% for the AIDS and 3.2% for the Rotterdam model. The correct model (the LES) is much more likely to find the trends when they are present in the data than when they are not -- the rejection frequency increases by 35%. The two incorrect models are only slightly more likely to reject when the trend is present (the frequencies of rejections for both the LA/AIDS and Rotterdam model increase by only about 3%).

These results raised concerns about whether the Rotterdam model is capable of detecting trends so we also tried using a Rotterdam model as the data generating mechanism. To do this we estimated a Rotterdam model, without trends or seasonality, using the Australian meat consumption data. We generated 100 data sets with this model by adding normal disturbances and tested for time trends using the Rotterdam and LA/AIDS models. As should be expected, the Rotterdam model found structural change about $100^*\alpha$ percent of the time, for a size α test, when there were no trends in the data generating mechanism: 8% of cases with a 5% test and 2% of cases with a 1% test. The LA/AIDS model found trends in 100% of cases with the 5% test and 99% of cases with a 1% test. This is a striking result: the LA/AIDS model was almost sure to find trends that weren't really there. Clearly the significant trends here are attributable entirely to specification error: the LA/AIDS did not do a good job of approximating the stable Rotterdam model.

Then we added trends to the simulation. The Rotterdam model detected trends much more often when they were present in the data (38% of cases using a 5% test; 18% using a 1% test). The LA/AIDS model actually found trends slightly less often when they were present using a 5% test (the rejection rate fell from 100% to 99%).

These results show that, when a demand system is estimated using data that were generated by another model, the probability of false structural change conclusions can increase or decrease. With data generated from the stable LES model, the tests with the LA/AIDS model are biased in favor of finding structural change and the tests with the Rotterdam model are biased against it. When the structural change is present in the data, the probability of finding it increases with the three models but much more so in the case of the true model, the LES. With data generated from the Rotterdam model, the LA/AIDS model found structural change all of the time, whether it was present or not. The Rotterdam model found trends about 38 percent of the time when they were in the data generating mechanism.

The Rotterdam and LES models performed very similarly: when each was the true functional form it found structural change in about α% of cases when it wasn't there and in about 40% of cases when it was there. The AIDS model falsely found structural change with a high frequency (27.6% for the LES data and 100% for the Rotterdam data) and that frequency did not increase much when the trend was present.

Thus, it seems that specification errors increase both the probability of false rejections and reduce the probability of true rejections. This is so, even when both the data generating mechanism and the estimated model are perfectly consistent with restrictions from the theory. These results also show that the use of flexible functional forms does not offer any protection from this misspecification bias.

Conclusion

We have not resolved whether Australian consumers have become more health conscious, nor whether their meat demand equations have changed for some other reason. Clearly variation in relative prices and total expenditures can account for much, if not all, of the changes in consumption patterns. Our nonparametric results support an assumption of stable, well-behaved preferences: relative prices and expenditures could account for all of the variation in per capita consumption.

The parametric results are mixed and we are not satisfied that any of our models is correctly specified. The Rotterdam model indicates there are not any trends; the Lewbel model indicates otherwise. The estimated trends were large enough to be economically very important; perhaps too large to be plausible. In the process of carrying out these tests we discovered that incorporating trends or seasonality in share equations, or in the Rotterdam model, is not easy; making sense of them is even harder.

The Monte Carlo work reinforces our previous conclusions derived with different data sets and different tests. Parametric tests for structural change (in this case tests for the presence of significant trends) are sensitive to specification errors. In particular, the LA/AIDS model tends to reject stable preferences far too often when it is not the correct functional form.

Thus we have resolved one thing: there are grounds for suspicion concerning the reliability of many of the previous studies of structural change in demand for meat. To what extent might the results be due to specification error? How should we proceed in order to reduce the risk of biased results due to arbitrary modeling choices? These are not new questions. The concluding comments of Booth and Judge (1956, p.583) seem as apt today as they were 35 years ago:

> *The research worker concerned with economic measurement needs more assurance that economic theory can allow for the random element in economic behavior, that the subjectivity of choice between alternative models can be reduced, and that accurate data can be made available. In this interdependent compromise between economic and statistical assumptions, manageability and realism, data and inference, there is little to warrant overconfidence in the estimates of economic parameters but much opportunity for econometric research.*

Footnotes

1. A fairly extensive listing of these studies is provided by Chalfant and Alston (1988) and by Moschini and Meilke (1989). Dahlgran (1988) summarizes the findings of a sample of them. Several are contained in the book edited by Buse (1989). Most of the studies have used U.S. data: Braschler (1983); Chalfant and Alston (1986, 1988); Chavas (1983); Choi and Soisin (1990); Dahlgran (1988); Haidacher (1983); Haidacher et al. (1982); Menkhaus et al. (1985); Moschini (1991); Moschini and Meilke (1984, 1989); Thurman (1987); Wohlgenant(1985). Studies using Canadian data include Alston and Chalfant (1991a); Atkins et al. (1989); Chen and Veeman (1989); and Young (1987). Previous studies using Australian data to test for structural change in meat demand include Chalfant and Alston (1986, 1988) and Martin and Porter (1985).

2. In this test, the data are checked for consistency with the Generalized Axiom of Revealed Preference (GARP). If the data are consistent with GARP, they *could have* been generated by a stable, well-behaved utility function. Canadian annual meat consumption data are also consistent with GARP (Alston and Chalfant, 1991a).

3. As pointed out by Chalfant and Alston (1988) -- following Landsburg (1981) and Varian (1982) -- when the nonparametric method is applied to data with strong trends in total consumption and relatively little price variation, budget lines will rarely cross and one is unlikely to find any violations of revealed preference axioms.

4. Recent studies that have estimated Australian meat demand equations include Alston and Chalfant (1987), Beggs (1988), Chalfant and Alston (1986), Fisher (1979), Main et al. (1976), Martin and Porter (1985), and Murray (1984). Most of these used single-equation models without imposing cross-equation restrictions derived from consumer theory. Fisher (1979) attempted a translog model and Chalfant and Alston (1986) attempted an LA/AIDS model.

5. The Martin and Porter data were also used by Alston and Chalfant (1987) and Chalfant and Alston (1988). As a departure from previous practice, we excluded mutton in response to concerns about the quality of the mutton data raised by Chalfant and Alston (1988) and others -- in particular Beggs (1989) noted that in later quarters the mutton prices had been inferred from lamb prices. In addition, fish is excluded due to unavailability of data (as is the usual practice with Australian meat demand).

6. Consistency with GARP is a necessary condition for weak separability of the group. This supports the exclusion of mutton and other goods from the analysis of within-group allocations of expenditure among the four meat types. This assumption is also supported to some extent by results from Alston and Chalfant (1987) and Chalfant and Alston (1988). LaFrance (1991) has raised some concerns about treating the constructed expenditure variable as exogenous in separable demand models. We acknowledge the correctness of his argument but choose to ignore the problem for the time being.

7. Moschini (1991) has made the same point.

8. For more detail on the AIDS model, see Deaton and Muellbauer (1980a, 1980b). For more detail on the translog, see Christensen, Jorgenson, and Lau (1975).

9. For more details on this model, see Theil (1971) and Theil and Clements (1987).

10. Deaton and Muellbauer (1980a) discuss the similarities and differences between the Rotterdam and AIDS models.

11. These maintained hypotheses could also be tested as a further specification check.

12. The LA/Lewbel rejected the LA/AIDS and the translog. However, we did not test the LA/Lewbel against the non-nested alternatives.

13. In a simple linear model with quantities (or logarithms of quantities) as the dependent variables a trend effect on quantity is constant each period. In a share equation the effects of an additive trend on quantities depend on the size of total expenditure and the good's price; when the trend enters nonlinearly, its effect on quantities depends on the other prices as well.

14. This work is closely connected to our previous work to evaluate implications of specification errors for Chow tests (Alston and Chalfant, 1991b).

15. Laitinen (1978) first showed that statistical tests of homogeneity were biased toward rejection, and several papers have since shown that the problem is pervasive and holds for other restrictions and for the various approaches to testing (Wald, likelihood ratio, and Lagrange multiplier tests); several of these studies are reviewed in Theil and Clements (1987).

16. Previously, we found an even higher rejection rate with a Chow test for a discrete structural change at the sample midpoint in an LA/AIDS model fit to data generated from the LES in this fashion: using a 5% test, the Chow test (falsely) rejected the stable model in 63% of cases, and using a 1% test it rejected in 38.6% of cases (Alston and Chalfant, 1991b).

17. Wald tests are more likely to reject than likelihood ratio tests (e.g., Berndt and Savin, 1977; Alston and Chalfant, 1991b).

18. The trend in the third share equation is obtained from adding up restrictions.

References

Alston, J.M., and J.A. Chalfant. "Weak Separability and a Test for the Specification of Income in Demand Models with an Application to the Demand for Meat in Australia." Australian Journal of Agricultural Economics 31(1987):1-15.

Alston, J.M. and J.A. Chalfant. "Can We Take the Con out of Meat Demand Studies." Western Journal of Agricultural Economics 16(1991a):(forthcoming).

Alston, J.M. and J.A. Chalfant. "Unstable Models from Incorrect Forms." American Journal of Agricultural Economics 73(1991b):(forthcoming).

Atkins, F.J., W. A. Kerr and D.B. McGivern. "A Note on Structural Change in Canadian Beef Demand." Canadian Journal of Agricultural Economics. 37(1989):513-524.

Beggs, J. "Diagnostic Testing in Applied Econometrics" Economic Record 64(1988):81-101.

Berndt, E.R., and N.E. Savin. "Conflict Among Criteria for Testing Hypotheses in the Multivariate Regression Model." Econometrica 45(1977):1263-1278.

Booth, E.J.R. and G.G. Judge. "The Impact of the Choice of Model on Measurements of Economic Behavior Relationships." Journal of Farm Economics. 38(1956):570-583.

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

Buse, R.C. (ed.) The Economics of Meat Demand. Proceedings of a Conference held October 20-21, 1986, Charleston, SC. Madison: Department of Agricultural Economics, University of Wisconsin, 1989.

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

Chalfant, J.A. and J.M. Alston. "Testing for structural Change in a System of demand for Meats in Australia. Australian Agricultural Economics Society Annual Conference, Canberra, February 1986.

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

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

Chen, P.Y. and M.M. Veeman "Demand for Meat in Canada: Model Specification and Structural Change." Staff Paper 89-11, Department of Rural Economy, University of Alberta, 1989.

Choi, S. and K. Sosin. "Testing for Structural Change: The Demand for Meat" American Journal of Agricultural Economics 72(1990):227-236.

Christensen, L.R., D.W. Jorgenson and L.J. Lau, "Transcendental Logarithmic Utility Functions." American Economic Review. 65(1975):367-383.

Dahlgran, R.A. "Changing Meat Demand Structure in the United States:Evidence from a Price Flexibility Analysis." North Central Journal of Agricultural Economics 10(1988):165-176.

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

Deaton, A.S. and J. Muellbauer. Economics and Consumer Behavior. Cambridge: Cambridge University Press, 1980b.

Gallant, A.R. "On the Bias in Flexible Functional Forms and an Essentially Unbiased Form: the Fourier Flexible Form." Journal of Econometrics. 15(1981):211-245.

Grunfeld, Y. "The Determinants of Corporate Investment." Unpublished Ph.D thesis. University of Chicago, 1958.

Haidacher, R.C. "Assessing Structural Change in the Demand for Food Commodities." Southern Journal of Agricultural Economics. 15(1983):31-37.

Kiefer, N. M., and J. G. MacKinnon. "Small Sample Properties of Demand System Estimates" In S. M. Goldfeld and R. E. Quandt, eds. Studies in Nonlinear Estimation New York: Ballenger? 1976.

LaFrance, J. "When is Expenditure 'Exogenous' in Separable Demand Models?" Western Journal of Agricultural Economics 16(1991):(forthcoming).

Laitinen, K. "Why is Demand Homogeneity so Often Rejected?" Economics Letters 1(1978):187-191.

Landsburg, S.E. "Taste Change in the United Kingdom, 1900-1955." Journal of Political Economy. 89(1981):92-104.

Lewbel, A. "Nesting the AIDS and Translog Demand Systems." International Economic Review. 30(1989):349-356.

McCloskey, D.N. The Rhetoric of Economics Madison: The University of Wisconsin Press, 1985.

Martin, W. and D. Porter. "Testing for Changes in the Structure of the Demand for Meat in Australia." Australian Journal of Agricultural Economics. 29(1985):16-31.

Menkhaus, D.J., J.S. St. Clair, and S. Hallingbye. "A Re-examination of Consumer Buying Behavior for Beef, Pork, and Chicken." Western Journal of Agricultural Economics. 10(1985):116-125.

Moschini, G. "Testing for Preference Change in Consumer Demand: An Indirectly Separable, Semiparametric Model." Journal of Business and Economic Statistics. 9(1991):111-117.

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

Moschini, G. and K.D. Meilke. "Modeling the Pattern of Structural Change in U.S. Meat Demand." American Journal of Agricultural Economics. 71(1989):253-261.

Murray, J. "Retail Demand for Meat in Australia: A Utility Theory Approach." Economic Record. 60(1984):45-56.

Stigler, G.J. The Theory of Price. New York:MacMillan Publishing Company Inc., 3rd Edition, 1966.

Theil, H. Principles of Econometrics. New York:John Wiley and Sons Inc., 1971.

Theil, H. and K.W. Clements Applied Demand Analysis: Results from System-Wide Approaches. Cambridge MA:Ballinger, 1987.

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

Varian, H. "The Nonparametric Approach to Demand Analysis." Econometrica 52(1982):945-973.

Varian, H. "Non-parametric Tests of Consumer Behavior." Review of Economic Studies 50(1983):99-110.

Wales, T.J. "A Note on Likelihood Ratio Tests of Functional Form and Structural Change in Demand Systems" Economics Letters 14(1984):213-220.

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

Young, L.J. Canadian Meat Demand. Working Paper 10/87. Ottowa: Agriculture Canada, Policy Branch, Commodity Coordination Directorate, June 1987.

Table 1: **Parameter Estimates from the Rotterdam Model (With and Without Trends), and Lewbel's Model and It's Special Cases (With Trends).**

	ROTTERDAM		LEWBEL		
	NO TREND	TREND	FULL	TRANSLOG	AIDS
α_0	---	---	7.4225	---	7.0884
α_{10}	---	---	2.5376	-0.5258	2.0262
α_{20}	---	---	-0.3501	0.4152	-0.1769
α_{30}	---	---	-1.7944	0.6352	-0.4263
Expenditure					
β_1	0.8653	0.8660	0.4273	---	0.3786
β_2	0.0865	0.0872	-0.1067	---	-0.0864
β_3	0.0387	0.0388	-0.3407	---	-0.1589
Prices					
γ_{11}	-0.2289	-0.2287	0.8351	-0.3461	0.6221
γ_{12}	0.1657	0.1663	-0.1308	0.0454	-0.0786
γ_{13}	0.0570	0.0569	-0.6990	-0.0016	-0.2706
γ_{14}	---	---	0.0543	0.0341	---
γ_{22}	-0.2333	-0.2326	-0.0289	-0.0469	-0.0437
γ_{23}	0.0568	0.0568	0.1240	0.0424	0.0630
γ_{24}	---	---	0.0236	0.0388	---
γ_{33}	-0.1247	-0.1235	0.4115	0.0296	0.0866
γ_{34}	---	---	-0.0034	0.0659	---
Trends					
τ_1	---	-7.08E-04	-1.24E-03	-1.02E-03	-1.23E-03
τ_2	---	-5.66E-04	-3.22E-04	-3.59E-04	-3.87E-04
τ_3	---	2.40E-04	8.10E-04	6.57E-04	7.84E-04
Seasonality					
θ_1	0.0534	0.0533	0.0503	0.0491	0.0553
θ_2	-0.0139	-0.0139	0.0395	0.0380	0.0430
θ_3	-0.0141	-0.0141	0.0244	0.0237	0.0261
θ_4	-0.0075	-0.0075	-0.0084	-0.0081	-0.0093
θ_5	0.0016	0.0016	-0.0078	-0.0072	-0.0085
θ_6	0.0101	0.0101	0.0030	0.0033	0.0030
θ_7	-0.0394	-0.0394	-0.0356	-0.0353	-0.0391
θ_8	0.0085	0.0085	-0.0285	-0.0280	-0.0309
θ_9	0.0034	0.0034	-0.0246	-0.0250	-0.0270

Table 2: **Likelihood Ratio (χ^2) Tests for Time Trends and Seasonality in the Lewbel, Translog, AIDS and Rotterdam Models**

TESTS	LEWBEL	TRANSLOG	ALMOST IDEAL	ROTTERDAM
		----- test statistics -----		
Trends				
H_0: $\tau_j = 0 \;\forall\; j$	120.96	121.98	126.12	2.52
H_0: $\tau_j = 0 \;\forall\; j$ (given $\theta_{ik} = 0 \;\forall\; i,k$)	111.88	93.74	95.90	3.06
Seasonality				
H_0: $\theta_{ik} = 0 \;\forall\; i,k$	146.98	160.14	164.42	195.06
H_0: $\theta_{ik} = 0 \;\forall\; i,k$ (given $\tau_j = 0 \;\forall\; j$)	137.90	131.90	134.20	195.60
Seasonality & Trends				
H_0: $\theta_{ik} = 0 \;\forall\; i,k$ and $\tau_j = 0 \;\forall\; j$	257.66	253.88	260.32	198.13

Notes: The numbers of parametric restrictions are 3 for trends, 9 for seasonality, and 12 for both trends and seasonality. The 95 percent critical values for the χ^2 with 3, 9, and 12 degrees of freedom are 7.81, 16.92, and 22.40, repectively. The 99 percent critical values for the χ^2 with 3, 9, and 12 degrees of freedom are 11.34, 21.7, and 27.7, repectively.

Table 3: **Likelihood Ratio (χ^2) Tests for Time Trends and Seasonality in the Approximate Lewbel, Translog and LA/AIDS Models**

TESTS	A/LEWBEL	TRANSLOG	LA/AIDS
		---- test statistics ----	
Trends			
H_0: $\tau_j = 0 \; \forall \; j$	139.90	121.98	113.26
H_0: $\tau_j = 0 \; \forall \; j$ (given $\theta_{ik} = 0 \; \forall \; i,k$)	98.48	93.74	89.08
Seasonality			
H_0: $\theta_{ik} = 0 \; \forall \; i,k$	154.04	160.14	155.08
H_0: $\theta_{ik} = 0 \; \forall \; i,k$ (given $\tau_j = 0 \; \forall \; j$)	112.62	131.90	130.90
Seasonality & Trends			
H_0: $\theta_{ik} = 0 \; \forall \; i,k$ and $\tau_j = 0 \; \forall \; j$	252.52	253.88	244.16

Notes: The numbers of parametric restrictions are 3 for trends, 9 for seasonality, and 12 for both trends and seasonality. The 95 percent critical values for the χ^2 with 3, 9, and 12 degrees of freedom are 7.81, 16.92, and 22.40, repectively. The 99 percent critical values for the χ^2 with 3, 9, and 12 degrees of freedom are 11.34, 21.7, and 27.7, repectively.

Table 4: Tests for Translog and AIDS Restrictions on Lewbel's Model

	RESTRICTION	
	TRANSLOG	ALMOST IDEAL
MODEL	H_0: $\beta_i = 0 \; \forall \; i.$	H_0: $\sum_{k=1}^{n} \gamma_{jk} = 0 \; \forall \; j.$
	------- test statistics -----	
TRUE LEWBEL		
Including Trends and Seasonality	11.08	4.98
Including Seasonality but No Trends	12.10	5.07
Including Trends but No Seasonality	24.24	22.42
No Trends and No Seasonality	7.30	7.64
LA/LEWBEL		
Including Trends and Seasonality	24.62	19.72
Including Seasonality but No Trends	6.70	12.80
Including Trends but No Seasonality	30.72	40.48
No Trends and No Seasonality	25.98	31.08

Notes: Both models require three parametric restrictions. The 95 percent critical value for the χ^2 with 3 freedom is 7.81 and the 99 percent critical value for the χ^2 with 3 degrees of freedom is 11.34.

Table 5: **Uncompensated Price and Expenditure Elasticities from Lewbel, Translog, AIDS and Rotterdam Models Including Trends and Seasonality**

	LEWBEL	TRANSLOG	AIDS	ROTTERDAM
Prices				
η_{bb}	-1.37	-1.34	-1.37	-1.30
η_{bl}	0.01	0.01	0.00	0.06
η_{bp}	-0.13	-0.15	-0.16	-0.20
η_{bc}	-0.20	-0.20	-0.19	-0.19
η_{lb}	0.70	0.68	0.69	0.77
η_{ll}	-1.41	-1.41	-1.40	-1.59
η_{lp}	0.10	0.15	0.13	0.26
η_{lc}	0.17	0.14	0.13	-0.01
η_{pb}	0.45	0.35	0.38	0.20
η_{pl}	0.14	0.17	0.16	0.27
η_{pp}	-0.94	-0.98	-0.95	-0.71
η_{pc}	0.23	0.26	0.27	0.03
η_{cb}	0.02	0.06	0.12	0.01
η_{cl}	0.29	0.26	0.26	0.07
η_{cp}	0.37	0.45	0.46	0.07
η_{cc}	-0.67	-0.72	-0.73	-0.21
Expenditure				
η_b	1.70	1.69	1.71	1.63
η_l	0.44	0.44	0.44	0.56
η_p	0.13	0.19	0.14	0.21
η_c	-0.02	-0.05	-0.11	0.07

Notes: The price elasticities are uncompensated (Marshallian) elasticities holding constant expenditure on the meat group: η_{ij} refers to the elasticity of quantity i with respect to price j and the subscripts refer to beef (b), lamb (l), pork (p) and poultry (c). The expenditure elasticities refer to the elasticity of quantity i with respect to total expenditure on meat. All elasticities are computed at every data point using predicted expenditure shares; the figures in the table are the means.

Table 6: **Uncompensated Price and Expenditure Elasticities from Lewbel, Translog, AIDS and Rotterdam Models Including Seasonality but Excluding Trends**

	LEWBEL	TRANSLOG	AIDS	ROTTERDAM
Prices				
η_{bb}	-1.46	-1.37	-1.32	-1.30
η_{bl}	-0.04	-0.02	-0.04	0.05
η_{bp}	0.04	0.01	-0.01	-0.20
η_{bc}	-0.01	-0.01	-0.03	-0.19
η_{lb}	0.56	0.62	0.60	0.77
η_{ll}	-1.49	-1.50	-1.53	-1.60
η_{lp}	0.32	0.37	0.40	0.26
η_{lc}	0.43	0.41	0.46	0.00
η_{pb}	0.59	0.41	0.37	0.20
η_{pl}	0.22	0.22	0.24	0.27
η_{pp}	-1.21	-1.20	-1.18	-0.71
η_{pc}	-0.11	-0.11	-0.10	0.03
η_{cb}	0.31	0.17	0.13	0.01
η_{cl}	0.47	0.40	0.46	0.08
η_{cp}	-0.22	-0.22	-0.21	0.08
η_{cc}	-1.31	-1.30	-1.32	-0.24
Expenditure				
η_b	1.47	1.39	1.40	1.63
η_l	0.17	0.09	0.07	0.56
η_p	0.51	0.68	0.67	0.21
η_c	0.74	0.95	0.95	0.08

Notes: The price elasticities are uncompensated (Marshallian) elasticities holding constant expenditure on the meat group: η_{ij} refers to the elasticity of quantity i with respect to price j and the subscripts refer to beef (b), lamb (l), pork (p) and poultry (c). The expenditure elasticities refer to the elasticity of quantity i with respect to total expenditure on meat. All elasticities are computed at every data point using predicted expenditure shares; the figures in the table are the means.

Table 7: **Importance of Trends in the Lewbel, Translog, AIDS, and Rotterdam Models – Average Quarterly Changes in Per Capita Consumption due to Trends (grams/quarter)**

	BEEF	LAMB	PORK	POULTRY
Trend Effects				
LEWBEL	-26.6	-10.1	15.4	29.8
TRANSLOG	-23.5	-10.3	14.2	28.4
AIDS	-24.4	-11.0	14.6	28.7
ROTTERDAM	-15.3	-15.2	4.0	39.2
Mean Consumption (Grams/Quarter)	11,567	4,204	3,724	4,447
Mean Budget Share	0.534	0.157	0.187	0.122

Note: The trends were not statistically significant in the Rotterdam model. The estimates are obtained for the Rotterdam model by taking the coefficient estimates that were statistically insignificant and treating them as parameters. Then, at the mean quantity and budget share, changes in quantity attributable to trend were calculated as:

$$\Delta q_i = \frac{\bar{q}_i \hat{\tau}_i}{\bar{S}_i}.$$