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Testing for Preference Change in Consumer Demand:

An Indirectly Separable, Semiparametric Model

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Abstract

This paper deals with the hypothesis that preference changes may partly explain observed consumption patterns. A model is developed based on indirect weak separability. The kernel estimator of the regression function models the unknown shape of demand functions, and a parametric structure models seasonality, dynamics, and preference change. This semiparametric specification is applied to U.S. meat demand. The results support the notion that changes in consumer preferences, due to consumers' awareness of the health hazards of cholesterol and saturated fat intake, may explain an increased consumption of white meat and a decreased consumption of red meat.

Key words: kernel estimation, nonparametric demand analysis, separability, U.S. meat demand.

1. INTRODUCTION

The analysis of U.S. meat demand has attracted considerable research attention in recent years, spurred by the steady increase in chicken consumption and the decline in beef consumption which have characterized demand patterns over the last two decades. Results based on aggregate data [see, among others, Chavas (1983), Moschini and Meilke (1989), and Thurman (1987)] show some evidence of structural change partly explaining the observed variation in consumption. This structural change may be explained in terms of changing preferences due to an increased concern among consumers about the health hazards of large intakes of cholesterol and saturated fats.

The conclusion that changing preferences may be responsible for variation in observed meat consumption has not gone unchallenged. Given that one cannot observe relevant features of the model generating the data (such as the shape of utility functions), or that certain features of the model are deliberately simplified for the sake of empirical tractability (for instance by aggregating across consumers), it is possible that evidence of structural change may reflect model misspecification of some kind. In particular, Chalfant and Alston (1988), Cox and Chavas (1988), and Wohlgenant (1985) emphasize the restrictive nature of parametric functional forms commonly used in modeling meat demand and the impact that this may have on a test for preference change.

To deal with the functional form issue in this context, this paper explores a new avenue that relies on the kernel estimator of the regression function (Ullah 1988). Specifically, a semiparametric model (Robinson 1988) is developed that allows for preference change and, under the null hypothesis of no preference change, does not make any restrictive

assumption about the shape of the true demand functions. Emphasis is on testing for preference change and on measuring the consumption bias effects of preference change in a manner similar to the analysis of technical change bias in production. Particular attention is devoted to the derivation of a separability assumption which results in estimable useful relationships.

This paper is organized as follows. The notion of separability is first reviewed, and it is shown that for the purpose of this study the assumption of indirect weak separability is more appealing than the commonly used direct weak separability. A semiparametric model of preference change is then specified, and the estimation procedure is reviewed briefly. In the empirical section this model is applied to U.S. meat demand by using quarterly data for the period 1967-1987. The results tend to support the contention that preference changes may have affected the composition of the U.S. meat consumption bundle.

2. SEPARABILITY AND PREFERENCE CHANGE

Separability assumptions are often invoked in empirical analysis to deal with the usual intractability of estimating a complete demand system [see Blackorby, Primont, and Russell (1978) for an extensive analysis of several separability concepts]. The most commonly used separability notion is that of direct weak separability (DWS) of the utility function. DWS can be formulated by writing the utility function as:

$$U = U(U^A(q^A), q^B) \quad (1)$$

where $U(\cdot)$ is a quasi-concave utility function, q^A is the subvector of

quantity consumed of the separable group A, q^B is the complement to q^A in the consumer bundle and includes all other goods, and the aggregator function $U^A(.)$ satisfies all the properties attributable to $U(.)$.

The interest in this specification lies in its consistency with the idea of two-stage budgeting. Indeed, DWS is necessary and sufficient for the second stage of two-stage budgeting, which allows the existence of conditional demand functions for the separable group of the type:

$$q_i = q_i(p^A, m^A) \quad i \in A \quad (2)$$

where q_i is the quantity demanded of good i , p^A is the vector of prices for the separable group, and m^A is the income allocated to this group.

Using the conditional demand functions in (2) to investigate the effect of preference changes for the nonhomothetic case, however, is not wholly appropriate. If changes in preferences affect the optimal value of m^A and the separable group is not homothetic, then changed preferences can affect consumption through the first stage income allocation effect. For example, if changed preferences entail a decrease in the allocation of income to the meat group, and this group is not homothetic with beef demand being more income elastic than poultry, then beef demand will decline relative to poultry even if the conditional demand functions as in (2) are not affected by structural change.

To specify a separable model suitable for the analysis of structural change this paper relies on the notion of indirect weak separability (IWS), which allows the indirect utility function to be written as:

$$V = V(V^A(p^A, m), p^B, m) \quad (3)$$

where m is total monetary income, and under standard regularity conditions $V(\cdot)$ is continuous, nonincreasing and quasi-convex in (V^A, p^B) , nondecreasing in m , and homogeneous of degree zero in (V^A, p^B, m) . The aggregator function $V^A(\cdot)$ is best interpreted as a price index for the separable group prices that also depends on the level of total income and has the same properties as the reciprocal of $V(\cdot)$. A characterization of IWS in terms of certain elasticities restrictions is reported in the Appendix, where it is shown that the restrictions on substitution possibilities implied by IWS are similar to, and not any more restrictive than, those implied by DWS.

It is known that the IWS in (3) is not implied by, nor does it imply, the DWS utility in (1) unless the aggregator function $U^A(\cdot)$ is homothetic. Because DWS is both necessary and sufficient for the second stage of consumer budgeting, it follows that IWS in general is not consistent with a meaningful definition of two-stage budgeting, which may explain why IWS typically is not used in empirical analysis [Caves, Christensen, and Herriges (1987), however, use IWS to test the homotheticity condition]. A useful simplification occurs, however, when one is interested in ratios of quantity demanded. Using Roy's identity, and considering the ratio q_i/q_j for $(i,j) \in A$, the IWS utility in (3) implies:

$$q_i/q_j = \frac{\partial V^A / \partial p_i}{\partial V^A / \partial p_j} = F^{ij}(p^A, m) \quad (4)$$

The important result here is that the ratios q_i/q_j of unconditional demand functions under IWS depend only on the subset of prices p^A and on total income. The corresponding degree of freedom gain, typical of second-

stage demand functions such as (2), in this case is attainable unconditionally as long as the variable of interest is given by quantity demand ratios. In analogy with the Hicksian notion of technical change, a preference change can be said to cause a consumption bias if it affects the ratios of quantity demanded. Hence, casting the analysis in terms of demand ratios, as implied by (4), is all that is needed to test for the existence of preference change effects and to characterize the nature of preference change. Consequently, the relationship in (4) provides a useful compromise between the advantages of a separability assumption and the desirability of unconditional results when testing for preference change.

3. A SEMIPARAMETRIC SPECIFICATION OF PREFERENCE CHANGE EFFECTS

Consistent with the objectives set out in the introduction, under IWS the problem is to model preference change effects without making unduly restrictive assumptions about $F^{ij}(p^A, m)$. To this end, this relationship is approximated by an amorphous function to be estimated by the kernel method, and the model is augmented by adding a parametric part that can account for preference changes, leading to a semiparametric specification. [Price-induced changes in tastes, as considered in Basmann, Molina, and Slottje (1983), can be captured by the nonparametric portion of the model.] The philosophy of this approach is similar to the "nonparametric" approach based on revealed preference theory developed by Varian (1982) and used in a context similar to ours by Chalfant and Alston (1988) and by Cox and Chavas (1988). The advantage of the procedure adopted here is that of allowing a straightforward statistical test of the hypothesis of preference change.

Let $Y^{ij} = \ln(q_i/q_j)$ represent the log of the demand ratio of interest, and let $\theta^{ij}(Z) = \ln[F_{ij}(p^A, m)]$, where $Z = p^A/m$ represent the normalized prices of the separable group so that homogeneity of degree zero of $F_{ij}(p^A, m)$ in (p^A, m) is satisfied. (The amorphous nature of the estimator of $\theta(Z)$ adopted does not allow a practical imposition of the Slutsky symmetry constraint. Failing to maintain this true restriction, however, will affect the efficiency of the model but not the validity of the results that follow.) Also, let X represent the variable indexing preference change. A semiparametric model of meat demand consistent with IWS, and which allows for the possible effects of changing preferences, can be written as:

$$Y_t^{ij} = \beta^{ij} X_t + \theta^{ij}(Z_t) + R_t^{ij} \quad (5)$$

where t indexes time series realizations, and R is a residual error term satisfying $E(R|X_t, Z_t) = 0$.

The introduction of the effects of preference change in the linear additive form of (5) can be interpreted in the variable addition framework of Pagan (1984) and allows a formal test of the hypothesis of absence of preference change, which reduces to testing $H_0: \beta^{ij} = 0$. Also, note that $\beta^{ij} = \partial \ln(q_i/q_j) / \partial X$ measures the rate of change in the demand ratio q_i/q_j associated with the variable representing preference change. Hence, this coefficient provides a measure of the consumption bias effects of preference change in a manner analogous to the notion of indirect Hicks neutrality/bias that can be used to characterize technical change in production (Lau 1978).

4. SEMIPARAMETRIC ESTIMATION

Following Robinson (1988), an estimator for β can be derived by noting that equation (5) implies:

$$Y_t - E(Y|Z_t) = \beta[X_t - E(X|Z_t)] + R_t \quad (6)$$

where the superscripted indices of (5) are dropped for notational convenience. To estimate β consistently it is crucial that the conditional expectations in (6) be evaluated without making unduly restrictive assumptions about $\theta(Z)$. This can be done by using the kernel method to estimate the conditional expectations in (6), and then estimating β by a no-intercept OLS given these conditional expectations. Robinson (1988) shows that the estimator of β derived in this fashion achieves root-N consistency and asymptotic normality in the case of independent observations, and suggests a number of extensions of this semiparametric model, including the case of stationary serially dependent data.

The kernel estimator of the conditional expectation is extensively reviewed in Bierens (1987) and Ullah (1988). Briefly, consider the problem of estimating the conditional expectation $E(Y|Z)$ at an arbitrary point from a sample of N realizations of the two random variates X and Z (Z can have more than one component). The sample $\{Y_1, Z_1; Y_2, Z_2; \dots; Y_N, Z_N\}$ is assumed to come from a continuous joint density function the value of which at the point (y, z) is given by $f(y, z)$. Because the conditional expectation of interest can be written as:

$$E(Y|Z=z) = \int_{-\infty}^{\infty} y \frac{f(y, z)}{f_1(z)} dy \quad (7)$$

where $f_1(z)$ is the marginal density of Z , one can obtain an estimate of $E(Y|Z=z)$ by estimating the density functions $f(y,z)$ and $f_1(z)$.

The kernel method offers an appealing density function estimation procedure (Silverman 1986). Using this estimator for the densities in (7) leads to a sample-based estimator of the conditional expectation of the type (Ullah 1988):

$$\hat{E}(Y|Z=z) = \int_{-\infty}^{\infty} y \frac{\hat{f}(y,z)}{\hat{f}_1(z)} dy = \sum_{i=1}^N Y_i r_i(z) \quad (8)$$

with:

$$r_i(z) = \frac{K[(z-Z_i)/h]}{\sum_{n=1}^N K[(z-Z_n)/h]} \quad (9)$$

where $K[.]$ is a multivariate kernel function satisfying $K(w) = \prod_{j=1}^M k(w_j)$ for $w_j = (z_j - Z_j)/h$, and h is the window width. Thus, the estimator $\hat{E}(Y|Z=z)$ is a weighted average of the sample's Y 's, where the weights depend on the evaluation point z , assigning the heaviest weight to the observations closest to the evaluation point.

To implement this nonparametric estimator, two main choices are required. First, a kernel function $K(w)$ must be chosen. As long as $K(w)$ satisfies certain regularity conditions, the choice of kernel has been shown not to be crucial (Epanechnikov 1969). A common practice, followed here, is to use the normal density:

$$k(w_j) = (2\pi)^{-1/2} \exp(-\frac{1}{2}w_j^2) \quad (10)$$

To deal with the asymptotic bias of the kernel estimator, Robinson (1988) suggests the use of higher order kernels, which can be obtained by linear combinations of (10), with the weights chosen such that the kernel has zero moments up to the desired order (see also Hong and Pagan 1988).

Second, one must choose the window width h . This is much more important because h controls the amount of smoothing imposed on the data. If h is too large the shape of the underlying density function is obscured (causing bias), whereas if h is too small spurious detail is introduced (causing imprecision). A viable trade-off between these two factors is achieved by using:

$$h = c N^{-1/(4+M)} \quad (11)$$

where c is a proportionality constant, N is the sample size, and M is the number of Z variables (nonparametric regressors) involved (Bierens 1987).

5. PREFERENCE CHANGE IN U.S. MEAT DEMAND

To apply the framework developed above to U.S. meat demand, it is postulated that an indirectly separable group for meat products comprises beef, pork, chicken, and fish. Thus, we consider six demand ratios: beef/pork, beef/chicken, beef/fish, pork/chicken, pork/fish, and chicken/fish. In accordance with the IWS assumption, the nonparametric portion of the model comprises four variables: the prices of beef, pork, chicken, and fish, all deflated by total per-capita income to ensure homogeneity.

In the parametric portion of the model a time trend represents the path of preference change. Under the assumption that preferences have changed

smoothly over the observation period, a trend variable can provide a useful proxy for these effects, and following Stone (1954), Deaton (1975), and Jorgenson and Lau (1975) this approach has been adopted in a number of studies. [The alternative of using the cholesterol information index reported by Brown and Schrader (1989), constructed from the number of articles in medical journals linking cholesterol in diet to heart disease, provided virtually the same results because of the high correlation of this index with trend. Thus, only the time trend models are reported].

The parametric portion of the model also allows for seasonality, in view of the quarterly data used, and both a static and a dynamic version of the semiparametric demand ratio model are reported. The normal kernel in (10) is used (higher order kernels gave very similar results, which are not reported here). The window width in (11) is used, with the constant c different for each variable (equal to the sample standard deviation). [Unreported analysis showed that replacing the constant 4 in (11) with 2 or with 6 gave very similar results, indicating some robustness of the semiparametric results to the choice of window width]. Estimation is carried out on a 386-based Zenith microcomputer utilizing user-written programs coded in GAUSS.

5.1 Data

Quarterly data for the period 1967 I to 1987 IV are used. Quantity demanded of beef and pork are given by per capita disappearance in retail weight as published by the U.S. Department of Agriculture in Livestock and Meat Situation (LMS) and in Livestock and Poultry Situation (LPS). The quantity of chicken is the total of young and mature chicken per capita

disappearance as published by the U.S. Department of Agriculture in Poultry and Eggs Situation (PES) and in LPS. Personal consumption expenditure for fish and seafood on a quarterly basis was obtained from unpublished U.S. Department of Commerce data; these data are expressed per capita by using estimates of U.S. population published in the Survey of Current Business (SCB) by the U.S. Department of Commerce. Retail prices for beef (choice) and pork as published in LMS and LPS, the retail price of frying chicken published in PES and LPS, and the CPI for fish and seafood from the U.S. Department of Commerce are used. Disposable income published in the SCB, divided by the population estimates also published in the SCB, provided the income term. Quantities are measured in lbs/capita for beef, pork, and chicken, and in 1987 dollar expenditures for fish. Prices, deflated by per-capita income, were normalized by their sample mean before utilization. The seasonal dummy variables are specified for the first three quarters. Time is measured with increments of 0.25 per quarter to ease the interpretation of the consumption bias coefficients.

5.2 A Static Model

The semiparametric model allowing for seasonality and preference change is written as:

$$Y_t^{ij} = \sum_{k=1}^3 \alpha_k^{ij} D_{kt} + \beta^{ij} t + \theta^{ij}(Z_t) + R_t^{ij} \quad (12)$$

where D_{kt} represent seasonal dummies. Some results from the estimation of this model are reported in Table 1. The first column reports the estimated β^{ij} , which gives a measure of the consumption bias attributable to preference changes. Because time trend was measured with increments of

0.25 per quarter, these coefficients are directly interpretable as annual rates of change in demand ratios attributable to preference changes (as such they satisfy $\beta^{ij} + \beta^{jk} = \beta^{ik}$). For example, the coefficient - 0.0252 for beef/chicken means that this demand ratio has declined by 2.52 % per year over the observation period for reasons not attributable to prices, income, and seasonal effects.

Five of the six estimated consumption bias measures are significantly different from zero at the 5 percent probability level, the only exception being the chicken/fish demand ratio. The sign of the estimated bias coefficients reveals the pattern of the hypothesized preference change. The change is biased against the two red meat products, beef and pork, relative to both white meat products, chicken and fish. The coefficients of the ratios involving chicken are among the most significant. The ratio of the two red meats shows a change in favor of beef and against pork, while no significant consumption bias is evident between the two white meats.

The second column of Table 1 reports the F-test of the hypothesis of no seasonal effects, i.e. $H_0: \alpha_1^{ij} = \alpha_2^{ij} = \alpha_3^{ij} = 0$, which is distributed with 3 and 80 degrees of freedom. Seasonality in meat consumption seems strong, in accordance with other studies using quarterly data. The R^2 's in Table 1 are computed as the ratio of explained variance to total variance of the dependent variable. With the exception of the beef/fish equation, they suggest a fairly good fit of the estimated models. The portmanteau Q statistics in the last column of Table 1 is based on the first 12 autocorrelations of the estimated residuals and is computed by using the amended formula of Ljung and Box (1978). This statistics, approximately

distributed as χ^2 with 12 degrees of freedom, is meant as a rough test of the independence of the residuals, which should be satisfied if the assumptions of the model are correct. It is clear that the Q statistics suggest some autocorrelation in the estimated residuals. The most likely reason for this outcome is the neglecting of the dynamic nature of the consumption data.

5.3 A Dynamic Model

Dynamic elements in observed consumption levels have been recognized in a number of studies, especially when frequent time series data are used (such as the quarterly data set of this study). This paper allows for dynamics by extending the parametric portion of the semiparametric model to include lagged values of the dependent variable. Although ad-hoc, this approach is consistent with the idea of habit persistence (Houthakker and Taylor 1970) and, as suggested in Robinson (1988), provides a viable introduction of dynamics in the semiparametric specification. Specifically, we consider lagged values of the dependent variable up to four quarters, so that the semiparametric model becomes:

$$Y_t^{ij} = \sum_{k=1}^3 \alpha_k^{ij} D_{kt} + \beta^{ij} t + \sum_{k=1}^4 \gamma_k^{ij} Y_{t-k}^{ij} + \theta^{ij}(Z_t) + R_t^{ij} \quad (13)$$

Some results from the estimation of this model are reported in Table 2. A feature of the dynamic specification chosen is that the coefficient of the time trend now gives the consumption bias of preference change conditional on the four lagged values of the demand ratio. Hence this coefficient has a short-run interpretation, and it is reported in the first column of Table 2. Because the lagged dependent variables are

different in each equation, the relationship $\beta^{ij} + \beta^{jk} - \beta^{ik}$ is no longer exactly satisfied. A long-run measure of consumption bias is attained as the effect embedded in the lagged dependent variables is taken into account, and the dynamic equation approaches the steady state solution. This long-run consumption bias is computed as $[\beta^{ij}/(1 - \sum_{k=1}^4 \gamma_k^{ij})]$, and the result is reported in the second column of Table 2 (the t-ratios for this column are computed from standard errors obtained from a linearization of this expression around the true values of the parameters).

The results shown in Table 2 are in general agreement with those in Table 1. The ratios of red meat products (beef and pork) to white meat products (chicken and fish) show a consumption bias against red meat and in favor of white meat. The ratio of red meat demands (beef to pork) and the ratio of white meat demands (chicken to fish) do not appear affected by preference change. Long-run consumption biases are larger than short-run ones, consistent with the fact that the estimated coefficients satisfy the stationarity condition for the dynamic specification used. The statistical significance of long-run and short-run measures are similar, with the exception of the beef/fish consumption bias which becomes insignificant in the long-run. [Dependence between observations, such as that associated with the presence of lagged dependent variables in (13), suggests caution in relying on the t-ratios discussed above because Robinson's (1988) asymptotic results do not apply to this case. However, unreported results from a bootstrapping experiment (Efron 1982) based on 1,000 replications support the accuracy of the asymptotic standard errors used for both the static and the dynamic models. In fact, the confidence bounds of the empirical distribution obtained from the bootstrap were

slightly tighter than those given by the standard errors used].

The estimated consumption bias measures are consistent with the hypothesis that preference change may have been caused by growing consumers' concern about intakes of cholesterol and saturated fats because this would conceivably shift demand away from products with high content of these nutrients (red meat products) toward products with a lower content of the same (white meat products). That the mix of red meat demand (beef to pork) and of white meat demand (chicken to fish) do not appear affected further supports this interpretation. These results are similar to the findings of Moschini and Meilke (1989), although their measure of consumption bias, estimated in a model that maintained direct weak separability and a parametric specification, did not consider pairwise comparison of individual meats.

The F-test of seasonality is similar to that of the static model and is distributed with 3 and 76 degrees of freedom. The F-test of dynamics concerns the hypothesis $H_0: \gamma_1^{ij} - \gamma_2^{ij} - \gamma_3^{ij} - \gamma_4^{ij} = 0$ and is distributed with 4 and 76 degrees of freedom. From Table 2 it emerges that seasonality and dynamics are important features of quarterly meat demand, a conclusion in agreement with many meat demand studies using quarterly data. This suggest an interesting comparison of the semiparametric model of this paper with the "nonparametric" method used by Chalfant and Alston (1988). Their results based on the quarterly Australian data set imply not only the absence of preference changes, but also the absence of seasonality and dynamic adjustment in consumption. The failure to detect at least seasonality may lend support to Varian's (1982) and Landsburg's (1981) observation that the nonparametric tests based on revealed preferences

may have little power with aggregate time series consumption data.

The R^2 's for the dynamic equations show an improved fit for all equations, especially for the equations involving fish. The Q statistics reported in Table 2 suggest that the dynamic elements can account for the serial correlation of the static equations. With the exception of the Q for the pork/chicken equation, which takes a borderline value, all of the Q statistics lead to a rejection of the hypothesis of serial dependence among the estimated residuals at the 5 percent probability level.

6. CONCLUDING REMARKS

This paper has shown that the assumption of indirect separability can be useful in testing for preference change. Unlike the use of second-stage demand functions from direct separability, indirect separability allows the estimation of consumption bias coefficients for unconditional demand functions, and therefore can provide a more appropriate test for the hypothesis of preference change. The semiparametric specification chosen retains the simplicity of a parametric form for the feature of interest, that is the effect of preference change, and it allows for a very flexible form for the unknown shape of the demand functions.

The empirical results suggest that U.S. meat demand is affected by a significant bias in consumption patterns over the period considered, in favor of white meat products (chicken and fish) and against red meat products (beef and pork). Of course, these results are conditional on the specification chosen. Although the indirectly separable model provides some generalization, it may still be too restrictive. Unfortunately, it is difficult to see how separability assumptions may be avoided in

specifying demand models to be estimated with time series data. What seems to emerge is that findings of preference change effects in meat demand cannot be ascribed entirely to the use of restrictive functional forms because the results of this paper were derived with minimal assumptions on the shape of demand functions under the null hypothesis of no preference change.

Although the estimated model can represent the direction, magnitude, and statistical significance of preference change effects, it cannot uncover the reasons for the existence of this change in meat demand. The results reported in this paper, however, are in agreement with what one would expect from a preference change spurred by consumers' growing appreciation of the health hazards of cholesterol and saturated fats in red meat products. If changing preferences are partly responsible for the evolution of observed U.S. meat consumption, as the results of this study suggest, then a policy response on the part of the red meat industry may be warranted, including possible adjustments towards the production of leaner meat and increased promotion and marketing efforts.

APPENDIX

Some interesting features of IWS are illustrated by the implied restrictions on substitution possibilities. If $(i,j) \in A$ and $k \notin A$, from (4) one obtains:

$$(\partial q_i / \partial p_k) - (\partial q_j / \partial p_k) F^{ij} \quad (A1)$$

Let S_{ik} denote the Slutsky substitution term such that $S_{ik}(p_k/q_i) = w_k \sigma_{ik}$, where σ denotes the Allen (1938) - Uzawa (1962) elasticity of substitution between the subscripted goods and $w_k = p_k q_k / m$ is the share of total income allocated to good k . Then using the Slutsky equation result $(\partial q_s / \partial p_r) = S_{sr} - q_r (\partial q_s / \partial m)$, and noting that $F^{ij} = q_i / q_j$, from (A1) it follows that IWS implies (for $(i,j) \in A$ and $k \notin A$):

$$\sigma_{ik} + \epsilon_j = \sigma_{jk} + \epsilon_i \quad (A2)$$

where ϵ denotes the income elasticity of the subscripted good.

This property of IWS can be contrasted with a similar implication of DWS. Blackorby and Russell (1976) showed that under homotheticity DWS implies $\sigma_{ik} = \sigma_{jk}$ (for $(i,j) \in A$ and $k \notin A$). Generalizing this result to the non-homothetic case of interest here yields:

$$\sigma_{ik} \epsilon_j = \sigma_{jk} \epsilon_i \quad (A3)$$

Hence, for both DWS and IWS the substitution possibilities of goods in the separable group with any one good outside the group are proportional to their income elasticities. Because $\sigma_{ik} > 0$ if goods i and k are net substitutes, the goods in the separable group with higher income elasticities will have greater substitution possibilities with any one good outside the group (that is, in both cases $\epsilon_i > \epsilon_j$ implies $\sigma_{ik} > \sigma_{jk}$). For DWS, if the goods in the separable group are all normal ($\epsilon_j, \epsilon_i > 0$), they will be either all net substitutes or all net complements with any

one good outside the group. On the other hand, for IWS the fact that the relationship is additive instead of multiplicative means that IWS does not require σ_{ik} and σ_{jk} to have the same sign, even if the i^{th} and j^{th} goods are normal (thus, IWS is somewhat less restrictive than DWS).

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TABLE 1 - Preference Change in Meat Demand: Static Model

Demand Ratio	Consumption Bias	F-test seasonality	R ²	Q(12)
Beef/Pork	0.0181 (3.4001)	31.21	0.91	43.34
Beef/Chicken	-0.0252 (-6.1122)	29.65	0.98	24.66
Beef/Fish	-0.0148 (-2.9030)	2.71	0.68	102.2
Pork/Chicken	-0.0433 (-8.5284)	84.45	0.96	44.69
Pork/Fish	-0.0329 (-5.2533)	18.80	0.89	75.57
Chicken/Fish	0.0104 (1.8753)	13.32	0.96	91.15

Note: Asymptotic t-statistics are reported in parenthesis.

TABLE 2 - Preference Change in Meat Demand: Dynamic Model

Demand Ratio	Consumption Bias short-run	Bias long-run	F-test seasonality	F-test dynamics	R ²	Q(12)
Beef/Pork	0.0066 (1.3231)	0.0112 (1.3360)	32.79	8.54	0.94	14.04
Beef/Chicken	-0.0187 (-4.6445)	-0.0309 (-4.3316)	12.94	6.14	0.98	11.26
Beef/Fish	-0.0095 (-2.6110)	-0.0518 (-1.4060)	7.67	19.86	0.84	15.33
Pork/Chicken	-0.0271 (-5.3302)	-0.0445 (-4.8292)	37.55	7.93	0.97	21.50
Pork/Fish	-0.0156 (-3.1543)	-0.0393 (-2.9160)	33.02	23.48	0.95	9.42
Chicken/Fish	0.0057 (1.5859)	0.0349 (1.3425)	17.06	29.06	0.98	8.45

Note: Asymptotic t-statistics are reported in parenthesis.