NON-NESTED TESTS AND AGRICULTURAL TRADE MODELS

BY

JUNG-HEE LEE

Bachelor of Economics
Chung-Ang University
Seoul, Korea
1985

Master of Science
Oklahoma State University
Stillwater, Oklahoma
1989

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Thesis Approved:

Wade Brorsen

Thesis Advisor

Dee Ahn

David Henneberry

Clement E. Ward

Thomas C. Collins

Dean of the Graduate College

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ESSAY I

A COX-TYPE NON-NESTED TEST FOR TIME SERIES MODELS

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ABSTRACT

This study introduces a Cox-type non-nested test that is a new approach to discriminating between linear or non-linear time series models. Based on the cox test of separate families of hypotheses, the Cox-type non-nested test uses Monte Carlo methods to obtain the distribution of Cox's non-nested test statistics. Using the maximum likelihood estimation technique, two competing time series models, generalized autoregressive conditional heteroschedasticity (GARCH) and exponential GARCH (EGARCH) models of daily spot prices of Deutsche Mark are estimated. Using Monte Carlo integration, then, the Cox-type non-nested test statistics for GARCH vs. EGARCH model are calculated. The EGARCH model is not rejected, while the GARCH model is rejected.

A COX-TYPE NON-NESTED TEST FOR TIME SERIES MODELS

1. Introduction

Cox (1961 and 1962) developed a likelihood ratio statistic to test separate families of hypotheses. Since then, numerous studies on tests of discriminating among separate models have been presented (for details, see Pereira, 1977). Econometricians have also adopted and studied the tests under the name of non-nested tests (Pesaran and Deaton; Fisher and McAleer; Dastoor). In time series studies, however, non-nested tests to discriminate among competing nonlinear time series models have not drawn much attention.

Since Engle (1982) introduced autoregressive conditional heteroskedasticity (ARCH) models, ARCH models have been developed in modified forms. The generalized ARCH (GARCH) model developed by Bollerslev (1986) has been the most widely used specification of ARCH. GARCH models impose restrictions on the parameters to assure positive variances. Nelson (1991) presented an alternative to GARCH models by changing GARCH to exponential GARCH (EGARCH). Unlike GARCH, EGARCH does not need inequality restrictions on parameters to assure a positive variance.

The objectives of this study are to develop a Cox-type non-nested test using Monte Carlo integration and use it to discriminate between two competing time series models, GARCH and EGARCH, in daily spot price of Deutsche Mark in terms of the United States dollar.

2. Cox's Test for Non-Nested Models

Suppose the observed value of a random vector $Y = (Y_1, \ldots Y_n)$ is to be used to test the null hypothesis, H_1, in which the probability density function (p.d.f.) of Y is $f(y, \delta_1)$, where δ_1 is an unknown vector parameter. As an alternative hypothesis, H_2 in which the p.d.f. of Y is $g(y, \delta_2)$ where δ_2 is an unknown vector parameter, is separate from H_1, that is, the two hypotheses are non-nested.

Cox (1961, 1962) suggested tests based on the log likelihood ratios

$$L_{12} = \log \frac{f(y, \hat{\delta}_1)}{g(y, \hat{\delta}_2)} = L_1(\hat{\delta}_1) - L_2(\hat{\delta}_2)$$

$$L_{21} = \log \frac{g(y, \hat{\delta}_2)}{f(y, \hat{\delta}_1)} = L_2(\hat{\delta}_2) - L_1(\hat{\delta}_1) \qquad (1)$$

where $L_1(\hat{\delta}_1)$ and $L_2(\hat{\delta}_2)$ are the maximum log-likelihood functions under H_1 and H_2, respectively. $\hat{\delta}_1$ and $\hat{\delta}_2$ are the maximum likelihood estimates of δ_1 under H_1 and δ_2 under H_2, respectively. Since the two hypotheses are non-nested, the choice of which is H_1 and which is H_2 is arbitrary.

To test H_1, the Cox statistic is

$$T_1 = L_{12} - E_1(L_{12}) \qquad (2)$$

where $E_1(L_{12})$ is the expected value of L_{12} under H_1. Cox (1961 and 1962) showed that under H_1, T_1 is asymptotically normally distributed with mean zero and variance V_1. The distribution of T_1 can be obtained analytically for many problems, but not time series.

2.1. Monte Carlo Implementation of Cox's Test

The expected value of L_{12} under H_1 is by definition

$$E_1(L_{12}) = \int L_{12}\, f(y, \delta_1)\, dy \qquad (3)$$

Similarly, since the mean of T_1 is zero the variance of T_1 under H_1 is by definition

$$V_1 = E_1[L_{12} - E_1(L_{12})]^2 \qquad (4)$$

Rewriting equation (4) yields

$$V_1 = \int [L_{12} - E_1(L_{12})]^2\, f(y, \delta_1)\, dy \qquad (5)$$

The integral of equations (3) and (5) can be evaluated by Monte Carlo methods when $\hat{\delta}_1$ is substituted for δ_1.

The ML estimators, $\hat{\delta}_1$ and $\hat{\delta}_2$, are consistent estimators of δ_1 and δ_2. By Slutsky's theorem and the dominated convergence theorem, then,

$$\text{plim}\left[\int L_{12}\, f(y,\hat{\delta}_1)\,dy\right] = \int L_{12}\, f(y,\delta_1)\,dy \tag{6}$$

Therefore, the Monte Carlo estimator of $E_1(L_{12})$ is consistent. By Slutsky's theorem and the dominated convergence theorem the Monte Carlo estimator of V_1 is also consistent:

$$\text{plim}\left[\int (L_{12} - E_1(L_{12}))^2 f(y,\hat{\delta}_1)\,dy\right] = \int [L_{12} - E_1(L_{12})]^2 f(y,\delta_1)\,dy \tag{7}$$

The final test statistic used is

$$NT_1 = \frac{\hat{T}_1}{\sqrt{\hat{V}_1}} \tag{8}$$

where $$\hat{T}_1 = \frac{L_{12} - \int L_{12}\, f(y,\hat{\delta}_1)\,dy}{\int \left(L_{12} - \int L_{12}\, f(y,\hat{\delta}_1)\,dy\right)^2 f(y,\hat{\delta}_1)\,dy}$$

is the Monte Carlo estimator of T_1 and $\hat{V}_1$ is the Monte Carlo estimator of V_1. Under H_1, $NT_1 \overset{d}{\rightarrow} N(0,1)$.

3. GARCH and EGARCH Models

Let ϵ_t denote a real-valued discrete-time stochastic process, and θ_{t-1} the set of all information available through time t-1. The GARCH(p,q) process is then given by

$$\epsilon_t | \theta_{t-1} \sim N(0, h_t) \tag{9}$$

where h_t is the conditional variance of ϵ_t. The GARCH(p,q) regression model is obtained by letting

$$h_t = \alpha_0 + \sum_i \alpha_i\, \epsilon^2_{t-i} + \sum_j \beta_j\, h_{t-j} \tag{10}$$

where $\alpha_0 > 0$, $\alpha_i >= 0$ for all i=1,...,p, $\beta_j >= 0$ for all j=1,...,q (restrictions can be relaxed somewhat for p>1 or q>1). On the other hand, the EGARCH(p,q) without the skewness term is written as (Hsieh,1991)

$$h_t = \exp[\alpha_0 + \sum_i \alpha_i \, |\epsilon_{t-i}| \, h_{t-i}^{-(1/2)} + \sum_j \beta_j \ln h_{t-j}] \tag{11}$$

The ϵ_t's may be innovations in a linear regression,

$$\epsilon_t = Y_t - X_t'b \tag{12}$$

where Y_t is the dependent variable, X_t is a vector of observations on explanatory variables including past realizations of Y_t, and b is a vector of unknown parameters to be estimated.[1] Then, the log likelihood function of a set of T observations is

$$L(\theta) = -\frac{T}{2}\ln 2\pi - \frac{1}{2}\sum_t \ln h_t - \frac{1}{2}\sum_t \frac{\epsilon_t^2}{h_t} \tag{13}$$

where $\theta = (\alpha_0, \alpha_i, \beta_j, b)$ and h_t is as defined in equation (10) for GARCH and as defined in equation (11) for EGARCH.

In general, GARCH(1,1) and EGARCH(1,1) models have been most widely used in the literature of ARCH type models because these are parsimonious. In this study, therefore, GARCH(1,1) and EGARCH(1,1) are used.

4. Empirical Results

The maximum likelihood parameter estimates and

[1]In this paper, X_t is a unit vector and thus b is the expected value of Y_t.

log-likelihood functions for GARCH and EGARCH models of daily spot prices of Deutsche Mark are presented in table 1. All estimates except the mean of Y_t in the GARCH(1,1) are statistically different from zero at the 5% significance level. Parameter estimates of GARCH(1,1) are similar to those obtained by Liu and Brorsen (1992) who used a different data source and a different algorithm.

Table 2 shows the results of Cox-type non-nested tests of GARCH and EGARCH using Monte Carlo integration. The calculated statistic, NT_1 under H_1: GARCH model is -1.69 which is smaller than $Z_{0.05} = -1.645$ based on a one-tailed test. Therefore, H_1: GARCH is rejected. In contrast, H_2: EGARCH model cannot be rejected because the calculated Cox statistic under H_2: EGARCH model, -0.68, is greater than $Z_{0.05} = -1.645$. Based on the Cox-type non-nested test, therefore, the EGARCH(1,1) is preferred to the GARCH(1,1) in modeling Deutsche Mark/U.S. dollar exchange rate.

5. Summary and Conclusions

This study introduced a Cox-type non-nested test that is a new approach to discriminating between linear or non-linear time series models. Based on the Cox test of separate families of hypotheses, our Cox-type non-nested test uses Monte Carlo integration to obtain the non-nested test statistics.

GARCH and EGARCH models of daily spot prices of Deutsche

Mark in terms of the United States dollar were estimated using maximum likelihood. The GARCH model was rejected, but the EGARCH model was not rejected. The results imply that the EGARCH models are preferred to GARCH models in modeling Deutsche Mark/dollar exchange rate. The Cox-type non-nested test procedures presented in this paper can be used to discriminate among competing linear or nonlinear time series models.

Appendix: Data and Estimation Procedure

This paper uses the daily spot price data of Deutsche Mark in terms of the United States dollar from January 1980 to September 1988. The data with 2212 observations are electronic data obtained from Technical Tools, an electronic data company. Since the daily spot prices are not stationary[2], they are transformed to the log percentage changes, Y_t, that is,

$$Y_t = \left(\ln \frac{PM_t}{PM_{t-1}} \right) 100$$

where PM_t is the daily spot price of Deutsche Mark at time t.

The widely used GARCH(1,1) and EGARCH(1,1) processes for h_t are adopted. The maximization of the log-likelihood functions for GARCH and EGARCH models is performed using the non-linear optimization algorithms of GAUSS (version 2.2). Steepest descent method with step length of one is first used and switched to Newton Raphson after 5 iterations.

First 20 observations out of 2212 observations are used to calculate an initial variance to use in the equation. First, the parameters, δ_1 and δ_2, and the log-likelihood functions,

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[2]To determine the stationarity of the daily spot prices, the price variable is transformed to natural log and Augmented Dickey-Fuller (ADF) test for a unit root is conducted. For the log price, the ADF test statistic is -1.11, while the critical value for rejecting the null hypothesis (the series has a unit root) is -2.57 at 10%. For the first differencing of log price, the ADF test statistic is -6.55, while the critical value at 10% is -2.57. These test results indicate that the price variable is not stationary, while the first differencing yields stationarity.

L_1 and L_2, of GARCH(1,1) and EGARCH(1,1) are estimated using maximum likelihood. Then using the ML estimates, $\hat{\delta}_1$ and $\hat{\delta}_2$, for GARCH and EGARCH models, Monte Carlo samples of 2264 observations are generated using the normal random number generator of GAUSS, "RNDN". For each sample, the first 52 observations are discarded to reduce effects of initial conditions.

Using 250 random samples based on the ML estimate of GARCH model, $\hat{\delta}_1$, the expected value and variance of L_{12} are obtained from

$$E_1(L_{12}) = \left[\sum_{i=1}^{n} (L_{1fi} - L_{1gi}) \right] \frac{1}{n}$$

$$V_1 = \left[\sum_{i=1}^{n} [L_{1fi} - L_{1gi} - E_1(L_{12})]^2 \right] \frac{1}{n-1}$$

where L_{1fi} is the maximum log-likelihood for the GARCH model from the i^{th} sample generated using $\hat{\delta}_1$, L_{1gi} is the maximum log-likelihood for EGARCH model from the i^{th} sample generated using $\hat{\delta}_1$, and n is the number of random samples generated. The standard deviation of L_{12} is obtained from $V_1^{1/2}$. Using fifty random samples based on the ML estimate of EGARCH model, $\hat{\delta}_2$, similarly, the expected value and the standard deviation of L_{21} are obtained.

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Table 1. Maximum Likelihood Estimates[a] of GARCH(1,1) and EGARCH(1,1) for Daily Deutsche Mark/Dollar Exchange Rate from Jan. 1980 to Sept. 1988

Parameters	GARCH	EGARCH
α_0	0.05461*	-0.28294*
α_1	0.15902*	0.30650*
β	0.76289*	0.90004*
b	-0.01804	-0.02605*
Log-likelihood	-2444.98	-2443.40

[a] An asterisk denotes rejection of the null hypothesis that the coefficient is zero at the 5 percent level of significance, using Wald-type test.

Table 2. Results[a] of Cox-type Non-nested Tests of GARCH(1,1) and EGARCH(1,1) for Daily Deutsche Mark/Dollar Exchange Rate from January 1980 to September 1988 using Monte Carlo Integration

Monte Carlo Estimates	GARCH	EGARCH
log-likelihoods	$E_1(L_{12}) = 4.24$	$E_2(L_{21}) = 3.51$
Variances	$V_1 = 11.81$	$V_2 = 7.95$
T	$T_1 = -5.82$	$T_2 = -1.97$
NT[b]	$NT_1 = -1.69^*$	$NT_2 = -0.68$

[a] An asterisk denotes statistical significance at the 5 percent level of significance.
[b] NT denotes Cox-type non-nested statistics having asymptotic standardized normal distributions.

ESSAY II

ALTERNATIVES TO THE ARMINGTON TRADE MODEL

ALTERNATIVES TO THE ARMINGTON TRADE MODEL

ABSTRACT

The Armington assumptions of homotheticity, weak separability, and single constant elasticity of substitution (CES) of import demands among import sources were tested and rejected using the double-log and AIDS models for U.S. beef import demand. An Orthodox non-nested test of the double-log and AIDS models was used to discriminate between the two alternatives to the Armington trade model. The non-nested tests failed to reject either the double-log model or the AIDS model. However, the estimated elasticities using the AIDS model were shown more plausible than those using the double-log model.

Key Words: Armington model, homotheticity, weak separability, constant elasticity of substitution (CES), non-nested test.

ALTERNATIVES TO THE ARMINGTON TRADE MODEL

INTRODUCTION

The Armington trade model differentiates commodity supply by country of origin. The Armington model assuming constant elasticity of substitution (CES) has been widely applied to agricultural import demand studies (e.g., modeling trade flows) where data limitations exist (Babula; Duffy et al.; Haniotis; Johnson et al.; Penson and Babula; Sarris). In modeling agricultural trade flows, the Armington model has been widely applied because of its parsimony with respect to parameters and its compatibility with demand theory (Alston et al. 1990).

However, the assumptions of homotheticity, weak separability, and single CES of import demands among import sources have prompted serious questions about the appropriateness of using the Armington model. The Armington assumptions have been rejected by previous studies which have tested the assumptions of Armington model using alternative models (Winters 1984; Alston et al. 1990; Ito et al. 1990). When the restrictions on demand are inappropriate, the parameters will be biased (Alston et al. 1990).

When the overall conclusions are that the Armington restrictions are inappropriate and, as a result, the parameters of Armington trade models are biased, then a question is what alternative model to the Armington model should be used to model agricultural trade flows and market shares using a less restrictive set of assumptions about demand relationships than Armington's?.

The objective of this paper is to determine 1) an alternative to the Armington trade model for U.S. source differentiated beef import demands using a non-nested test and 2) elasticities of U.S. source differentiated beef import demands. We also test the Armington restrictions of homotheticity, separability, and single CES of import demands among import sources using both of the Almost Ideal Demand System (AIDS) specification of import demand and the double-log specification of import demand in which the Armington model is nested, for the U.S. beef import demands.

Historically, the United States has been one of the largest beef importers as well as one of the largest exporters in the world (Table 1). U.S. beef imports took about 19% of the world beef imports in value over 1970-1990, and 13% in 1990. However, most previous U.S. beef trade studies have not paid attention to the import side.

The Armington assumptions of homotheticity and weak separability of import demands among import sources were tested with the AIDS model using United Kingdom import data for manufactured goods by Winters, and with the AIDS and the double-log model using cotton and wheat trade data by Alston et al. Ito et al.tested the assumptions of homotheticity and single CES using rice trade data. They all rejected the Armington restrictions.

Winters suggested the AIDS model as an alternative to the Armington model. Alston et al. also presented the double-log model and the AIDS model as possible alternatives to the Armington model. However, their studies were not concerned about the alternatives because they simply focused on testing the Armington assumptions. Ito et al. introduced an alternative to the Armington model, a "modified Armington model". The model is expressed in double-log form. The Armington model is nested in the double-log form. However, the AIDS model is non-nested in the double-log form.

To model trade flows, two alternative models introduced by Alston et al. and Winters were recently used by Haden and Honma. Haden used the AIDS model to estimate demand elasticities for cigarettes disaggregated by Japanese, U.S. and rest-of-world sources for the Japanese market. The double-log import demand model was used by Honma to analyze growth in Japan's horticultural trade with developing countries. However, they used the models without considering other alternatives to the Armington model.

None of the above studies has tested all three assumptions of the Armington model (i.e., homotheticity, weak separability, and single CES). No attempts to discriminate among alternative models (e.g., non-nested tests) have been tried. Therefore, this study goes beyond past work in critiquing the Armington model and evaluating alternatives to it.

ARMINGTON TRADE MODEL

Armington uses a two stage budgeting procedure. In the first stage, the importing country's expenditure allocation among imported goods is determined by maximizing utility subject to a budget constraint. In the second stage, Armington assumes the utility is weakly separable among n goods (i.e., beef is weakly separable from pork). Using weak separability among n goods, total expenditure on each good is allocated among m different kinds of products that are differentiated by origin. This allocation is determined as minimizing the cost of purchasing total imports of a good.

Armington also introduced the assumptions that (a) elasticities of substitution in each market are constant and (b) the elasticity of substitution between any two products competing in a market is the same as that between any other pair of products competing in the same market. The assumptions of (a) and (b) imply a single CES (i.e., only one CES among import demands for m different products in a market).

Under the assumption of weak separability among n goods, using the CES within-group specification, the Armington model has the following form:

$$(1) \quad X_{ij} = b_{ij}^{\sigma i} X_i (P_{ij}/P_i)^{-\sigma i}$$

where X_{ij} is the imported good i from source j, b_{ij} is a constant, σ_i is the elasticity of substitution in the i^{th} good market, and P_i is the import price index depending only on the within-group prices, expressed as $P_i = \Sigma_k (X_{ik}/X_i) P_{ik}$. Taking the logarithms on both sides of equation (1) yields:

(2) $\log X_{ij} = \sigma_i \log b_{ij} + \log(E_i/P_i) - \sigma_i \log(P_{ij}/P_i)$

where E_i is total expenditure on imports of good i from all sources (i.e., $Ei = P_i X_i$).

Equation (1) can also be written as follows:

(3) $W_{ij} = b_{ij}^{\sigma i} (P_{ij}/P_i)^{1-\sigma i}$

where W_{ij} is $(P_{ij}X_{ij})/(P_iX_i)$ expressing the expenditure share of imports for good i from source j. Taking the logarithms on both sides of equation (3) yields:

(4) $\log W_{ij} = \sigma_i \log b_{ij} + (1-\sigma_i) \log(P_{ij}/P_i)$

The CES specification (4) implies the assumptions of weak separability among m different products (i.e., Canadian beef is weakly separable from Australian beef in U.S. beef import) and homotheticity of import demands for m different products. Homotheticity implies that the expenditure share of imports for good i from source j is independent of the budget allocated to the imports for good i in an importing country. In addition, the coefficients for the relative price terms (i.e., $1-\sigma_i$) are the same for all products because of the assumption of a single CES.

DOUBLE-LOG IMPORT DEMAND MODEL

Consider the following double-log specification[3] of the within-group allocation of expenditures among m sources of imports of good i within a single importing country (Alston et al.):

(5) $\log X_{ij} = \alpha_{ij} + \Sigma_k \Gamma_{ijk} \log(P_{ik}/P^*) + \beta_{ij} \log(E_i/P^*)$

where β_{ij} denotes the expenditure elasticity for good i from source j, Γ_{ijk} is the compensated cross-price elasticity of good i from source k on the demand of good i from source j, $E_i = \Sigma_j P_{ij} X_{ij}$ for all j where P_{ij} and X_{ij} are price of good i from source j and quantity of good i from source j, respectively, $\log P^* = \Sigma_j W_{ij} \log(P_{ij})$ which is called "Stone's price index". The Marshallian price elasticities are obtained from the Slutsky equation $\Gamma m_{ijk} = \Gamma_{ijk} - \beta_{ij} W_{ik}$ at the mean shares. However, Alston et al treated Γ_{ijk} in equation (5) as the Marshallian price elasticity which should be the Hicksian price elasticity.

The double-log model is homogeneous of degree zero in all prices and total expenditure for good i. However, the

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[3] This specification is the Hicksian demand function. This specification is originated from the Marshallian demand function; $\log X_{ij} = \alpha_{ij} + \Sigma_k \Gamma^m_{ijk} \log P_{ik} + \beta_{ij} \log E_i$, in which coefficients for prices are the Marshallian price elasticities. Using the Slutsky equation; $\Gamma^m_{ijk} = \Gamma_{ijk} - \beta_{ij} W_{ik}$, the Marshallian equation is transformed to the Hicksian demand function which is homogeneous of degree zero in prices and expenditure, at least approximately. (see Deaton and Muellbauer(1980b:p61-62) for details)

theoretical restrictions of symmetry and adding-up in consumer behavior cannot be globally imposed on the double-log specification of the demand system (Deaton and Muellbauer 1980b). The Armington restrictions of homotheticity, weak separability, and single CES of import demands among m sources will be tested on the double-log import model, separately and all together.

The alternative to the Armington model presented by Ito et al. is also expressed in double-log forms. A modified Armington model by Ito et al. is as follows:

(6) $\log(X_{ij}/X^{*}_{i}) = a_{ij} + \beta_{ij}\log(E_i/P^{*}) + \sigma_{ij}\log(P_{ij}/P_i)$

where X^{*}_{i} is total import demand of the i^{th} good. The modified model still retains one of the Armington assumptions (i.e., weak separability of import demands among m different sources). Therefore, the restriction of weak separability may be tested on the double-log import model.

To estimate the double-log import model of equation (5), Seemingly Unrelated Regression (SUR) estimators are used. For separability, H_0: $\Gamma_{ijk} = 0$ for all $k \neq j$ (weak separability means that only the own price and group price are included in the model). For homotheticity, H_0: $\beta ij = 1$ for all j (homothetic demands imply that in the absence of price changes import budget share will not change). For the single CES, H_0: $\Gamma_{ijj} = \Gamma_{ikk} = -\sigma_i$ for all j, k. The Wald F-test will be used.

ALMOST IDEAL DEMAND SYSTEM (AIDS) MODEL SPECIFICATION

In the AIDS model of import demand (Deaton and Muellbauer 1980a), the budget share of imports for good i from source j:

(7) $W_{ij} = \mu_{ij} + \Sigma_k \pi_{ijk} \log P_{ik} + \Omega_{ij} \log(E_i/P^*)$, $j = 1,\ldots,m$,

where P_{ij} is import price of good i from source j, E_i is total expenditure on imports of good i from all sources, and $\log P^* = \Sigma_k W_{ik} \log(P_{ik})$. When Stone's price index is used in the AIDS model, it causes a simultaneity problem because the dependent variable expenditure share is used to calculate $\log P^*$. To avoid the problem, We follow Eales and Unnevehr and use the lagged share to calculate $\log P^*$ (i.e., $\log P^* = \Sigma_k W_{ikt-1} \log(P_{ikt})$. In estimating demand systems, expenditure is also not exogenous because expenditures are used to compute the dependent variable (Attfield, 1985, and LaFrance, 1991). Expenditure being correlated with the error terms causes estimates to be biased and inconsistent. Most past literature just assumes the simultaneity to be small and then ignores the problem. We follow Blundell (1987) and use the Wu-Hausman test[4] to determine if expenditure can be treated as exogenous.

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[4] Let v_{i1} error term of the AIDS model, equation 7. For the purpose of testing the exogeneity assumption of expenditure the equation for $\log(E_i/P^*)$ is approximated by $\log(E_i/P^*)_t = a_i + \Sigma_j f_{ij} \log P_{ijt} + g_i \log(E_i/P^*)_{t-1} + h_i Y_t + v_{i2t}$ where t is time, Y is a country's total income, and v_{i2} is a random error term. The error term v_{i1} is partitioned as follows $v_{i1} = \zeta v_{i2} + e_i$ where ζ is correlation parameters such that $E(v_{i2}, e_i) = 0$ and therefore e_i is independent of v_{i2}. The residual v_{i2} is included in equation 7. An F test for the inclusion of the residuals provides an asymptotically efficient exogeneity test.

The adding-up condition is satisfied with $\Sigma_j\ \mu_{ij} = 1$, $\Sigma_j\ \pi_{ijk} = 0$ and $\Sigma_j\ \Omega_{ij} = 0$, while homogeneity and symmetry are satisfied with $\Sigma_k\ \pi_{ijk} = 0$ and $\pi_{ijk} = \pi_{ikj}$, respectively. SUR estimators are used to estimate the AIDS import demand models of equation (7) with symmetry and homogeneity imposed. Since the adding-up condition causes the contemporaneous covariance matrix to be singular, an equation in the model is arbitrarily deleted and then others in the model are estimated by Seemingly Unrelated Regression (SUR).

For homotheticity, H_0: $\Omega_{ij} = 1$ for all j. For the single CES, H_0: $\pi_{ijj} = \pi_{ikk}$ for all j, k. For a test of separability (between import sources), Winters's test method will be used. Alston et al. also followed Winters and tested whether the price from a particular import source contributes anything to the otherwise complete allocation model. This condition is a necessary consequence of separability (Alston et al.).

To test if one source is separable from the others, for each import source a reduced AIDS model excluding it is estimated. Then it is tested whether the price of excluded import source has any influence on the import shares of import sources included in the reduced AIDS model. Now, the reduced AIDS model is specified as follows:

(8) $W_{ij} = \mu_{ij} + \Sigma_{k\neq h}\ \pi_{ijk}\log P_{ik} + \Omega_{ij}\log(E_i/P^*) + \delta_{ijh}\log P_{ih}$,

for all j, j ≠ h, and all h, where P_{ih} is import price of good i from source h. The reduced AIDS model has m-1 import sources since an import source (h) is excluded from the full AIDS model. Symmetry and homogeneity are imposed on the reduced models of equation (8). To estimate the reduced models, an equation in each reduced model is deleted for the singular contemporaneous covariance matrix and then SUR estimators are used.

For separability between an import source h and other import sources among m-1 import sources, H_0: $\delta_{ijh} = 0 \; \forall \; j, \; j \neq h$. Homotheticity and single CES are tested as well. H_0: $\Omega_{ij} = 0 \; \forall \; j, \; j \neq h$ for homotheticity, while H_0: $\pi_{ijj} = \pi_{ikk} \; \forall \; j, k, \; j \neq h, \; k \neq h$ for σ, single CES.

NON-NESTED TESTS

In this study, the double-log import model is expressed linearly in double-log form and the AIDS import model is expressed linearly in semi-log form, so that one model cannot be obtained from the other by simply imposing restrictions on the parameters. Hence, the models are non-nested.

The dependent variable in the double-log specification (2) is the imported quantity, while in the AIDS specification (7) it is the import share. Before testing the non-nested models, the double-log specification and the AIDS specification, the double-log specification (2) is converted as follows (for more details, see Appendix):

$$(9) \quad \log W_{ij} = \alpha_{ij} + \Sigma_{k \neq j} \Gamma_{ijk} \log P_{ik} + (\Gamma_{ijj}+1) \log P_{ij} - (\Sigma_k \Gamma_{ijk}+1) \log P^* + (\beta_{ij}-1) \log(E_i/P^*)$$

Let $\theta_{ij} = - \Sigma_k \Gamma_{ijk}$. Then rewriting equation (9) yields:

$$(10) \quad \log W_{ij} = \alpha_{ij} + \Sigma_{k\neq j} \Gamma_{ijk}\log P_{ik} + (\Gamma_{ijj} + 1)\log P_{ij} + (\theta_{ij} - 1)\log P^* + (\beta_{ij} - 1)\log(E_i/P^*)$$

Using the AIDS import demand and the double-log import demand in share form (10), the non-nested hypothesis is tested to discriminate between the two models:

$$(11) \quad H_0: W_{ij} = \mu_{ij} + \Sigma_k \pi_{ijk}\log P_{ik} + \Omega_{ij}\log(E_i/P^*)$$

$$(12) \quad H_1: \log W_{ij} = \alpha_{ij} + \Sigma_{k\neq j} \Gamma_{ijk}\log P_{ik} + (\Gamma_{ijj}+1)\log P_{ij} + (\theta_{ij}-1)\log P^* + (\beta_{ij} - 1)\log(E_i/P^*)$$

To test the non-nested hypothesis, an orthodox non-nested test will be used, which nests the two non-nested models in a more general model.

$$(13) \quad W: W_{ij} = c_{ij} + \Sigma_k \Phi_{ijk}\log P_{ik} + s_{ij}\log P^* + b_{ij}\log(E_i/P^*)$$

In this study, however, the non-nested models have dependent variables in a different form; one is W_{ij} (i.e., the AIDS model is a semi-log specification) and the other is $\log W_{ij}$. In this case, the Box-Cox transformation can be used (Judge et al.). Using the Box-Cox transformation function, the general model is specified as:

$$(13) \quad (W^{\lambda}_{ij} -1) / \lambda = c_{ij} + \Sigma_k \Phi_{ijk}\log P_{ik} + s_{ij}\log P^* + b_{ij}\log(E_i/P^*)$$

where λ is the Box-Cox transformation parameter. When $\lambda = 1$, equation (13) represents a semilogarithmic model. When $\lambda = 0$, equation (13) represents a double-log model. the value of λ in the general model is constrained to be between 0 and 1.

Using the Box-Cox transformation, the general model in equation (13) nests the AIDS and double-log import models. For all three models, homogeneity conditions are imposed: $\Sigma_k \Phi_{ijk} = 0$ for AIDS model, $\Sigma_k \Phi_{ijk} + s_{ij} = 0$ for the general model and the double-log model. To test AIDS, the null hypothesis is H_0: $s_{ij} = 0$ and $\lambda = 1$ for all j. For the double-log model, the null hypothesis is H_0: $\lambda = 0$.

The two non-nested models are now nested in the general model above. Each non-nested model is tested against the general model using the Chi-square likelihood ratio test for which the test statistic is

(14) $2[L_u - L_r] \overset{d}{\rightarrow} \chi^2(J)$,

where L_u is the log-likelihood for the unrestricted model (i.e., general model) and L_r is the log-likelihood for the restricted model with J restrictions imposed.

DATA

Major beef exporters to the U.S. in terms of value were Australia, New Zealand, and Canada for 1970-1991 (Table 2). These three countries accounted for more than 90% of the total value of U.S. beef imports over the period. Therefore, beef exporters to the United States are divided into Australia, New Zealand, Canada, and the rest of the world (ROW) in this paper.

Annual time series data are used for U.S. beef imports from Canada, Australia, New Zealand, and ROW from 1970 to 1991.

U.S. import data for fresh or frozen beef are obtained from Foreign Agricultural Trade of the United States (FATUS), USDA/ERS. U.S. import unit values of beef are used as proxies for import prices.

EMPIRICAL RESULTS

The estimated results of the Armington and double-log import models are summarized in Table 3. The Armington model is estimated by imposing the Armington restrictions on the double-log import model. System weighted R^2 is 0.93 for the Armington model and 0.98 for the double-log model. The estimated elasticities of substitution for the Armington model are the same for all equations due to the single CES assumption. The elasticity of substitution is 1.22 and significantly different from zero at the 1-percent significance level. Ordinarily, it would be expected that σ exceeds unity: an improvement in competitiveness should yield an increased share (Armington). The estimated elasticity of substitution implied from equation (4) suggests that a relative fall in P_{ij} yields an increase in the market share of X_{ij}. Therefore, the Armington model seems to be plausible. However, all Armington restrictions are rejected (Table 4). Separability, homotheticity, and a single CES of U.S. beef import demands among four major import sources are rejected at the 10-percent significance level by separate tests as well as a complete system test at the 1-percent significance level. The rejection of separability among import sources within group implies the generalized Armington model with separability by Ito et al. is also rejected.

In the Wu-Hausman tests for the AIDS model and the double-log model in share form, the F values were 0.59 with 3 and 45 degrees of freedom and 1.83 with 4 and 52 degrees of freedom, respectively. Thus the Wu-Hausman test cannot reject the hypothesis in which expenditure in the AIDS model and the double-log model are not correlated with the error terms. This test results imply that expenditures in the models in this study can be treated as exogenous.

The estimated results of the double-log import model are presented in Table 3. System weighted R^2 is 0.98 for the double-log model. In the double-log model, estimates for all Marshallian own price elasticities are negative as expected except for New Zealand. Estimated own price elasticities of Canadian beef and ROW beef appear very elastic to own prices. All estimates for expenditure elasticities are positive as expected. The elasticities for Australia and New Zealand are significant at the 1-percent significance level. The estimated expenditure elasticities indicate that Australia will have relatively more share of U.S. beef imports than other beef exporters to U.S when U.S. expenditures for beef imports increase.

Table 5 presents the estimated AIDS model for U.S. beef import demands. System weighted R^2 is 0.37. Marshallian elasticities are in Table 6. All own price elasticities are negative. The price elasticities are -2.93, -1.19, and -2.22 and significant at the 5-percent significance level for Canada, Australia, and ROW, respectively, while the elasticity for New Zealand's beef is insignificant at the 5-percent significance level. These results indicate elastic U.S. import demands for beef from Canada, Australia, and ROW.

Positive cross price elasticities indicate competitive relations, while negative cross price elasticities indicate complementary relations. Estimated cross price elasticities indicate that Canadian beef competes with ROW beef in the U.S. beef market. This positive elasticity is significant at the 5-percent significance level. All other cross price elasticities are insignificant.

Expenditure elasticities for all countries are positive and significant at the 5-percent significance level except for ROW, which is insignificant at the 5-percent significance level. The estimated expenditure elasticities are almost unitary elastic (1.1) for Canadian beef, elastic (1.25 and 1.43) for Australian beef and New Zealand beef, respectively. These results indicate that in U.S. beef imports, the Canadian import share will be relatively constant with increasing U.S. expenditures for beef imports, while Australian and New Zealand import shares will increase. On the other hand, if the United States cuts the budget of imported beef, the Australian and New Zealand import shares will decrease.

AIDS model F-test results are summarized in Tables 7 and 8. Using the AIDS model, homotheticity is rejected by complete system tests at the 1-percent significance level, while a single CES is not rejected (Table 7). In the joint test, homotheticity and single CES are rejected at the 5-percent significance level. Separability is rejected at the 1-percent significance level in Canadian, Australian, and New Zealand cases (Table 8). On the other hand, homotheticity and single CES are not rejected at the 5-percent significance level. By the joint tests for separability and homotheticity, separability and homotheticity are rejected at the 1-percent significance level in Canadian, Australian, and New Zealand cases. Using the joint tests for all Armington restrictions, separability, homotheticity, and single CES are rejected at the 1-percent significance level in Canadian, Australian, and New Zealand cases.

The Armington model is based on the joint assumptions of separability, homotheticity, and single CES. Therefore, the joint tests are critical to the Armington model. Overall, the Armington restrictions of separability, homotheticity, and single CES are rejected using the AIDS and the double-log import models.

Table 9 presents the non-nested test results. Both the AIDS model and the double-log import model are not rejected at the 5% significance level in all equations. In general, therefore, either model seems appropriate for modeling source differentiated U.S. beef import demands.

SUMMARY AND CONCLUSIONS

By the joint and individual tests for the Armington restrictions using the double-log import model, all Armington restrictions were rejected. Using the AIDS model, the joint tests for the Armington model also rejected the Armington restrictions. Alston et al. already rejected the restrictions using world cotton and wheat trade data. Therefore, these results imply that the Armington restrictions are inappropriate for modeling agricultural import demands and cause specification errors by omitting relevant explanatory variables (e.g., import prices from competing sources within group).

The AIDS and double-log models with less restrictive assumptions than Armington's have been considered as possible alternatives to the Armington model (Winters and Alston et al.). This paper tested the non-nested models, the AIDS and double-log models of source differentiated U.S. beef import demands. The non-nested tests showed that both the double-log import model and the AIDS model cannot be rejected. However, the estimated elasticities using the AIDS model were shown more plausible than those using the double-log model. In addition, the AIDS model permits imposing the theoretical properties of demand, while the double-log model only allows homogeneity. Therefore, the AIDS is preferred to the double-log model in modeling source differentiated U.S. beef import demands.

All estimates of own price elasticities in the AIDS import model are negative. The results show elastic U.S. import demands for beef from Canada, Australia, and ROW. Estimated expenditure elasticities for all countries in the AIDS import model are positive. Cross price elasticities indicate that Canadian beef competes with beef from ROW in the U.S. beef market. The estimates of expenditure elasticities indicate that in U.S. beef imports, Canadian import share will be relatively constant with increasing the U.S. expenditures for beef imports, while Australian and New Zealand import shares will increase. On the other hand, if the United States cuts the budget of imported beef, Australian and New Zealand import shares will decrease.

The non-nested tests in this paper provide a general approach to discriminate between the AIDS and double-log models.

Table 1. Trends in the Value Share of World Beef Imports; 1970-1990

Year	Major Importers		
	EC*	U.S.	Japan
1970	0.53	0.31	0.01
1975	0.58	0.16	0.02
1980	0.50	0.18	0.05
1985	0.49	0.15	0.07
1990	0.49	0.13	0.13
Average	0.52	0.19	0.06

* EC denotes the European Community as EC-12 unit throughout.

Source: FAO Trade Yearbook, Food and Agriculture Organization (FAO) of the United Nation, various issues.

Table 2. Summary Statistics and Trends for U.S. Beef Import Expenditure shares; 1970-1991

	Canada	Australia	New Zealand	ROW[a]
Mean	0.082	0.49	0.26	0.17
Std. Dev.	0.034	0.035	0.044	0.063
Minimum	0.022	0.43	0.20	0.097
Maximum	0.14	0.56	0.35	0.27
1970	0.07	0.45	0.21	0.26
1975	0.02	0.53	0.21	0.23
1980	0.06	0.56	0.23	0.15
1985	0.14	0.44	0.30	0.12
1991	0.11	0.48	0.31	0.10

[a] ROW is the rest of the world.

Source: Foreign Agricultural Trade of the United States (FATUS), ERS/USDA, various issues.

Table 3. Elasticities[a] of U.S. Beef Import Demand using Armington and Double-log Models

Import Source(j)	Armington	Double-Log				
		Prices				Expenditure
	σ	Canada	Australia	New Zealand	ROW	
Canada	1.22*	-3.48*	-0.80	13.55*	3.88	0.03
	(5.45)	(-2.24)	(-0.11)	(3.21)	(1.43)	(0.06)
Australia	1.22*	0.10	-0.94	-1.84*	-0.25	1.43*
	(5.45)	(0.28)	(-0.62)	(-2.05)	(-0.42)	(11.67)
New Zealand	1.22*	-0.29	3.88	3.47*	2.39*	1.02*
	(5.45)	(-0.49)	(1.43)	(2.15)	(2.29)	(4.66)
ROW[b]	1.22*	0.22	-9.72*	-10.73*	-7.32*	0.27
	(5.45)	(0.22)	(-2.19)	(-4.06)	(-4.29)	(0.75)
System[c] R^2	0.93	0.98				

[a] The Marshallian price elasticities of the double-log import demand model are obtained from $\Gamma^m_{ijk} = \Gamma_{ijk} - \beta_{ij} w_{ik}$ at the mean shares where Γ denotes the Hicksian price elasticities and β. * denotes statistical significance at the 5 percent level of significance. t-values are in parentheses.

[b] ROW is the rest of the world.

[c] This is system weighted R^2.

Table 4. F-Test Results of the Double-log Model

	S[a]	H[b]	Single CES	S-H-Single CES[c]
F-values[d]	5.57**	3.26**	2.03*	7.82**
d.f.	(12,60)	(4,60)	(3,60)	(19,60)

[a] S is separability among import sources.
[b] H is homotheticity.
[c] S-H-single CES denotes separability, homotheticity, and single CES: All restrictions are tested at once.
[d] ** denotes statistical significance at the 5 percent level of significance, while * at the 10 percent level of significance.

Table 5. SUR Estimates of the AIDS Model for U.S. Beef Import Demand

Import Sources (j)	SUR Estimates[a]				
	π_1	π_2	π_3	π_4	Ω_j
Canada	-0.16**	0.029	-0.062	0.19**	0.008
	(-3.42)	(0.46)	(-1.03)	(3.75)	(0.20)
Australia	0.029	-0.033	0.021	-0.017	0.12**
	(0.46)	(-0.15)	(0.11)	(-0.22)	(2.18)
New Zealand	-0.062	0.021	-0.035	0.076	0.11*
	(-1.03)	(0.11)	(-0.18)	(0.86)	(1.71)
ROWb	0.19**	-0.017	0.076	-0.25**	-0.24**
	(3.75)	(-0.22)	(0.86)	(-2.77)	(-3.93)
System[c] R^2			0.37		

[a] ** denotes statistical significance at the 5 percent level of significance, while * at the 10 percent level of significance. t-values are in parentheses.
[b] ROW is the rest of the world.
[c] This is system weighted R^2.

Table 6. Marshallian Elasticities[a] of U.S. Beef Import Demand using of the AIDS model

Import Source(j)	Prices				
	Canada	Australia	New Zealand	ROW	Expenditure
Canada	-2.93*	0.30	-0.78	2.31*	1.10*
	(-5.14)	(0.38)	(-1.06)	(3.73)	(2.20)
Australia	0.038	-1.19*	-0.022	-0.077	1.25*
	(0.29)	(-2.60)	(0.06)	(-0.51)	(10.9)
New Zealand	-0.28	-0.13	-1.25	0.22	1.43*
	(-1.16)	(-0.16)	(-1.63)	(0.65)	(5.70)
ROW[b]	1.23*	0.60	0.80	-2.22*	-0.41
	(4.0)	(1.27)	(1.54)	(-4.22)	(-1.14)

[a] In AIDS model the uncompensated (i.g., Marshallian) price elasticities are given by

$\epsilon_{ij} = - \delta_{ij} + (\pi_{ij}/w_j) - \Omega_i(w_j/w_i)$,

where δ_{ij} is unity it i=j and zero otherwise. The expenditure elasticity is given by

$\eta_i = 1 + (\Omega_i/w_i)$.

t-values are in parentheses. * denotes statistical significance at the 5 percent level of significance.

[b] ROW is the rest of the world.

Table 7. F-test Results[a] of the AIDS Model

	Homotheticity	Single CES	H-Single CES[b]
F-values	4.79**	0.11	2.87**
d.f.	(3,51)	(2,51)	(5,51)

[a] These are from complete system tests. ** denotes statistical significance at the 5 percent level of significance.
[b] H-Single CES denotes homotheticity and single CES: Homohteticity and single CES are tested at once.

Table 8. AIDS Model F-test Results[a] using Winters's Method

Separable Countries (h)	S^b $\delta_{ijh}=0$ d.f. (2,33)	H^c $\Omega_{ij}=0$ d.f. (2,33)	Single CES $\pi_{ijj}=\pi_{ikk}$ d.f. (1,33)	$S\text{-}H^d$ d.f. (4,33)	$S\text{-}H\text{-}CES^e$ d.f. (5,33)
Canada	14.71**	2.47*	1.24	9.88**	8.63**
Australia	6.56**	3.43*	0.10	9.525**	9.03**
New Zealand	9.11**	2.13	0.002	7.78**	6.52**
ROW[f]	0.44	0.28	0.13	0.47	0.48

[a] ** denotes statistical significance at the 5 percent level of significance, while * at the 10 percent level of significance.
[b] S is separability.
[c] H is homotheticity.
[d] S-H denotes separability and homotheticity: Separability and homotheticity are tested at once.
[e] S-H-CES denotes separability, homotheticity, and single CES: All restrictions are tested at once.
[f] ROW is the rest of the world.

Table 9. Non-nested Model χ^2 Test Results

Import Sources	AIDS[a] d.f.= 2	Double-log[b] d.f. = 1
Canada	1.84	2.70
Australia	1.29	0.26
New Zealand	3.42	0.0
ROW[c]	2.74	0.46

[a] These statistics refer to $s_{ij}=0$ and $\lambda = 1$ where λ denotes Box-Cox transformation parameter on W_{ij}. Thus, the degree of freedom for the Chi-square equals 2.

[b] These statistics refer to $\lambda = 0$. Thus, the degree of freedom for the Chi-square is 1.

[c] ROW is the rest of the world.

APPENDIX

Derivation of the double-log model with dependent variable as import quantity to one as import share:

$$\text{(A1)}\quad \log X_{ij} = \alpha_{ij} + \sum_k \Gamma_{ijk} \log\frac{P_{ik}}{P^*} + \beta_{ij} \log\frac{E_i}{P^*}$$

$$= \alpha_{ij} + \Gamma_{ij1} \log\frac{P_{i1}}{P^*} + \ldots + \Gamma_{ijm} \log\frac{P_{im}}{P^*} + \beta_{ij} \log\frac{E_i}{P^*}$$

$$= \log e^{\alpha_{ij}} + \log\prod_k^m \left(\frac{P_{ik}}{P^*}\right)^{\Gamma_{ijk}} + \log\left(\frac{E_i}{P^*}\right)^{\beta_{ij}}$$

Rewriting equation (A1),

$$\text{(A2)}\quad \log X_{ij} = \log\left[e^{\alpha_{ij}} \prod_k^m \left(\frac{P_{ik}}{P^*}\right)^{\Gamma_{ijk}} \left(\frac{E_i}{P^*}\right)^{\beta_{ij}} \right]$$

Eliminating logarithms of both side of equation (A2),

$$\text{(A3)}\quad X_{ij} = e^{\alpha_{ij}} \prod_k^m \left(\frac{P_{ik}}{P^*}\right)^{\Gamma_{ijk}} \left(\frac{E_i}{P^*}\right)^{\beta_{ij}}$$

Multiplying equation (A3) by (P_{ij}/E_i),

$$\text{(A4)}\quad W_{ij} = \left(\frac{P_{ij}}{E_i}\right) e^{\alpha_{ij}} \prod_k^m \left(\frac{P_{ik}}{P^*}\right)^{\Gamma_{ijk}} \left(\frac{E_i}{P^*}\right)^{\beta_{ij}}$$

Taking logarithms on both side of equation (A4),

$$\text{(A5)}\quad \log W_{ij} = \alpha_{ij} + \log\frac{P_{ij}}{E_i} + \sum_k \Gamma_{ijk} \log\frac{P_{ik}}{P^*} + \beta_{ij} \log\frac{E_i}{P^*}$$

Rewriting equation (A5),

$$\text{(A6)}\quad \log W_{ij} = \alpha_{ij} + \sum_{k \neq j} \Gamma_{ijk} \log P_{ik} + (\Gamma_{ijj} + 1)\log P_{ij}$$

$$- (\sum_k \Gamma_{ijk} + 1)\log P^* + (\beta_{ij} - 1)\log(\frac{E_i}{P^*})$$

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ESSAY III

SOURCE DIFFERENTIATED U.S. BEEF DEMAND AND SEPARABILITY

SOURCE DIFFERENTIATED U.S. BEEF DEMAND AND SEPARABILITY

ABSTRACT

All previous studies on source differentiated import demand have assumed separability between domestic and imported goods. Separability is a strong assumption because a decision to import is often based on changes in domestic demand. Using the AIDS, dynamic AIDS, and Rotterdam models, the separability test results suggest that U.S. beef import demand differentiated by source is not independent of domestic beef demand, which can be estimated independently of other domestic meat. The Rotterdam model was used for U.S. beef demand differentiated by source because it was the best fit for the data. The empirical results indicate that U.S. demand for nonfed beef is highly elastic, while U.S. demand for fed beef is inelastic. The results also indicate that with increasing U.S. total expenditure on beef, nonfed beef will have a relatively higher value share in the U.S. beef market. Among source differentiated imported beef, U.S. import demand for ROW's beef is inelastic.

Key Words: source differentiated beef demand, AIDS, dynamic AIDS, Rotterdam model, multi-stage budgeting, separability.

SOURCE DIFFERENTIATED U.S. BEEF DEMAND AND SEPARABILITY

INTRODUCTION

The United States is simultaneously one of the largest beef exporters and importers in the world (Alston et al. 1989; Lee et al. 1992). Overall, the United States is a net beef importer.

Most U.S. beef exports are grain fed beef which can be differentiated from grass fed beef in terms of quality and use (Eales and Unnevehr; Brester). The majority of U.S. beef imports are grass fed beef. The major beef exporters to the United States have been Australia, New Zealand, and Canada. Imported beef from Australia and New Zealand is grass fed, while imported beef from Canada is grain fed. Imported beef from Australia and New Zealand accounted for about 75 percent of the total value of U.S. beef imports over 1970-1991 (Lee et al. 1992). The imported grass fed beef is mostly used for the school lunch program and hamburger. USDA data (FATUS, USDA) shows that unit values per metric ton for export beef are much higher than unit values for import beef. The higher unit values support the idea that the United States exports higher quality fed beef and imports lower quality grass fed beef.

Recently, Japan, the largest importer of U.S. beef, agreed to release restrictions on beef imports. There are two major opinions on the benefits from liberalization. The United States expects that this liberalization will increase U.S. beef exports to Japan and benefit U.S. beef producers (Coyle and Dyck, 1989). However, a question is whether U.S. exports will grow faster under the liberalized policy and whether producer benefits will outweigh U.S. consumer costs (Alston et al., 1989). They argue

> even if the U.S. beef industry were to gain from Japanese liberalization, net U.S. welfare may decline. If Japanese liberalization increases world beef trade, the price of beef should increase in every country which is closely linked to world prices. Thus, U.S. beef producers could indirectly gain from Japanese beef trade liberalization, even if U.S. beef exports to Japan decline. However, since the U.S. is likely to remain a net importer, its consumers should lose more from a beef price increase than its producers gain.

The net benefits from beef trade liberalization are not clear a priori. To determine whether or not producers would gain more than consumers lose, an overall study including beef exports, beef imports, and domestic beef demand and supply should be conducted. However, most previous U.S. beef trade studies have not focused on the import side. Therefore, this study focuses on U.S. beef imports and is expected to provide a partial answer to this debate.

It is now important to determine the relationship between imported beef and domestic beef in the U.S. market. This can be accomplished by measuring the effects on expenditure shares among source differentiated imported beef, domestic beef, and other meat when prices of imported beef and domestic beef, and expenditures change. To measure the effects of the price and expenditure changes on expenditure shares, the demand elasticities of source differentiated import beef and domestic beef are estimated. If imported beef is assumed to be differentiated from domestic beef, import beef and domestic beef are treated as separated products in the model.

Armington originally developed a trade model to differentiate import demands by source so that disaggregated import demands differentiated by source can be estimated instead of aggregate import demands. However, the Armington model is based on strong assumptions of homotheticity, separability among different sources, and single constant elasticity of substitution (CES). These assumptions have been rejected by hypothesis tests (Winters; Alston et al.; Seale et al.; Yang et al.; Lee et al.).

As alternatives to the Armington trade model, the almost ideal demand system (AIDS) and the double-log model[5] (Winters; Alston et al.; Haden; Honma; Ito et al.; Yang and Koo; Lee et al., 1992) have been widely used. More recently, the Rotterdam model is also widely used for agricultural trade (Lee et al.; Sparks et al.; Weatherspoon and Seale; Brester).

All previous studies on source differentiated import demands have assumed separability between domestic and

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[5] The double-log model can hold only one of the theoretical demand properties, homogeneity, while all the restrictions can be imposed and tested on the AIDS and the Rotterdam models. The study results by Lee et al. (1992) suggested that the AIDS model fits U.S. beef import data better than the double-log model.

imported goods. This separability implies that the marginal utility of an imported good depends only on the consumption of other imported goods so that demands for imported goods are estimated conditional on total import expenditure and independently of demand for domestic goods. However, separability is a strong assumption because a decision to import is often based on changes in domestic demand.

The objectives of this paper are to determine 1) the U.S. budget allocation on beef including domestic beef and source differentiated imported beef, and 2) the elasticities of a source differentiated U.S. beef demand system using the AIDS, the dynamic AIDS, and the Rotterdam models. The assumption of separability between domestic vs. imported beef, imported beef vs. other imported meat, and imported beef vs. other domestic meat are tested in this study.

METHODS

Multi-stage Budgeting Theory

Suppose that a country is treated as an individual consumer. In multi-stage budgeting (Figure 1), a country allocates expenditures first between food and non-food and next between meat and non-meat. In the third stage, then, a country allocates the expenditures on meat, E, among beef, pork, poultry, and other meat differentiated by source. Following Armington, meat differentiated by kind is called "good" (e.g., beef, pork, poultry, or other meat) and a good differentiated by source is called "product" (e.g., Canadian beef, Australian beef, etc.).

In this paper, unlike the literature on source differentiated import demands, block independence between domestic and imported goods is not assumed. Therefore, the marginal utility of an imported good depends not only the consumption of other imported goods but also the consumption of domestic goods. In multi-stage budgeting, a necessary and sufficient condition for the second and third stages is weak separability of the direct utility function over broad groups of goods (Deaton and Muellbauer, 1980b). The third stage is estimated using the demand system approach such as the AIDS, the dynamic AIDS, and the Rotterdam models.

Almost Ideal Demand System (AIDS) Model

The AIDS specification of demand (Deaton and Muellbauer 1980a) is derived by specifying an expenditure function representing the Price-Independent-Generalized-Logarithmic (PIGLOG) class of preferences. The homogeneous expenditure function in prices, p, is defined as

(1) $\ln E(u,p) = a(p) + ub(p)$

where u is a country's utility. The functional forms for a(p) and b(p) are chosen such that the first and second derivatives of the expenditure function can be set equal to those of an arbitrary expenditure function, thus satisfying the necessary condition for flexibility of functional form. These forms are written as

$$(2) \quad a(p) = \alpha_0 + \sum_i \sum_h \alpha_{i_h} \ln P_{i_h} + \frac{1}{2} \sum_i \sum_j \sum_h \sum_k \gamma^*_{i_h j_k} \ln P_{i_h} \ln P_{j_k}$$

where i and h indicate a good and a source, respectively, p_{i_h} is the price of good i from source h, p_{j_k} is the price of good j from source h, and

$$(3) \quad b(P) = \beta_0 \prod_i \prod_h P_{i_h}^{\beta_{i_h}}$$

so that the AIDS expenditure function is written

$$(4) \quad \ln E(u,p) = \alpha_0 + \sum_i \sum_h \alpha_{i_h} \ln p_{i_h} + \frac{1}{2} \sum_i \sum_h \sum_j \sum_k \gamma^*_{i_h j_k} \ln p_{i_h} \ln p_{j_k} + \beta_0 \, u \prod_i \prod_h p_{i_h}^{\beta_{i_h}}$$

By Shephard's lemma, the price derivatives of the expenditure function are the quantities demanded: $\frac{\partial E(u,p)}{\partial p_{i_h}} = q_{i_h}$.

Multiplying by $\frac{p_{i_h}}{E(u,p)}$ yields

$$(5) \quad \frac{\partial \ln E(u,p)}{\partial \ln p_{i_h}} = \frac{p_{i_h} q_{i_h}}{E(u,p)} = w_{i_h}$$

where w_{i_h} is the expenditure share of good i from source h in total consumption of a food group. Thus, the differentiation of equation (3) with respect to $\ln p_{i_h}$ yields the expenditure shares as a function of prices and utility as

(6) $$w_{i_h} = \alpha_{i_h} + \sum_j\sum_k \gamma_{i_h j_k} \ln p_{j_k} + \beta_{i_h} u \prod_i \prod_h p_{i_h}^{\beta_{i_h}}$$

where

(7) $$\gamma_{i_h j_k} = \frac{1}{2}(\gamma^*_{i_h j_k} + \gamma^*_{j_k i_h}) = \gamma_{j_k i_h}$$

Solving equation (4) with respect to u and substituting this into equation (6) gives the AIDS model in expenditure share form

(8) $$w_{i_h} = \alpha_{i_h} + \sum_j\sum_k \gamma_{i_h j_k} \ln p_{j_k} + \beta_{i_h} \ln \frac{E}{P^*}$$

where P^* is price index defined by follows

(9) $$\ln P^* = \alpha_0 + \sum_i\sum_h \alpha_{i_h} \ln p_{i_h} + \frac{1}{2}\sum_i\sum_h\sum_j\sum_k \gamma_{i_h j_k} \ln p_{i_h} \ln p_{j_k}$$

Equation (8) is a first-order approximation to the general unknown relation between W_{i_h} , lnE, and lnp's. However, the system is nonlinear due to the nonlinear price index equation. To allow for linear estimations, equation (9) is replaced with Stone's price index to which P^* is approximately proportional:

(10) $$\ln P^* = \sum_i\sum_h w_{i_h} \ln p_{i_h}$$

Equation (8) using Stone's price index is called the linear approximate AIDS (LA/AIDS) model (Blanciforti and Green). It combines the best of the theoretical features of both the Rotterdam and translog models with the ease of estimation of the Linear Expenditure System (LES). However, using the Stone's price index causes a simultaneity problem because the dependent variable expenditure share appears on the right hand side of equation (8). To avoid simultaneity, the lagged share is used to calculate $\ln P^*$ (Eales and Unnevehr, 1988).

The theoretical restrictions for adding-up, homogeneity, and Slutsky symmetry, respectively, require

$$(11)\quad \sum_i\sum_h \alpha_{i_h} = 1,\ \sum_i\sum_h \gamma_{i_h j_k} = 0,\ \sum_i\sum_h \beta_{i_h} = 0$$

$$(12)\quad \sum_j\sum_k \gamma_{i_h j_k} = 0$$

$$(13)\quad \gamma_{i_h j_k} = \gamma_{j_k i_h}$$

The conditions for homogeneity and symmetry follow from the homogeneity of the expenditure function (1), while the condition for adding-up follows from equation (7). Unrestricted estimation of the system (8) will automatically satisfy the adding-up restrictions. The restrictions of homogeneity and symmetry can be imposed and tested on equation (8). With restrictions (12) and (13), therefore, the estimated AIDS equations add up to a given total group expenditure, are homogeneous of degree zero in prices and the total expenditure, and satisfy the Slutsky symmetry requirement (Deaton and Muellbauer, 1980b).

Uncompensated price elasticities of the AIDS model are

$$(14)\quad \epsilon_{i_h j_k} = -\delta_{i_h j_k} + \frac{\gamma_{i_h j_k}}{w_{i_h}} - \beta_{i_h}\left(\frac{w_{j_k}}{w_{i_h}}\right)$$

where $\delta = 1$ *for* $i=j$ *and* $h=k$, otherwise zero, and the average expenditure shares are used for the expenditure shares. The variance of uncompensated price elasticities is

$$V(\epsilon_{i_h j_k}) = V(\delta_{i_h j_k}) + (\frac{1}{w_{i_h}})^2 V(\gamma_{i_h j_k}) + (\frac{w_{j_k}}{w_{i_h}})^2 V(\beta_{i_h})$$
$$- \frac{2}{w_{i_h}} Cov(\delta,\gamma) + 2(\frac{w_{j_k}}{w_{i_h}}) V(\delta,\beta)$$
$$- 2(\frac{w_{j_k}}{w_{i_h}^2}) V(\gamma,\beta)$$

The compensated price elasticities are

(15) $$\epsilon^*_{i_h j_k} = \epsilon_{i_h j_k} + w_{j_k} + \beta_{i_h}(\frac{w_{j_k}}{w_{i_h}}) = -\delta_{i_h j_k} + \frac{\gamma_{i_h j_k}}{w_{i_h}} + w_{j_k}$$

The variance of compensated price elasticities is

$$V(\epsilon^*_{i_h j_k}) = V(\delta_{i_h j_k}) + \frac{1}{w_{i_h}^2} V(\gamma_{i_h j_k}) - \frac{2}{w_{i_h}} Cov(\delta,\gamma)$$

The expenditure elasticity is

(16) $$\eta_{i_h} = 1 + \frac{\beta_{i_h}}{w_{i_h}}$$

The variance of expenditure elasticities is

$$V(\eta_{i_h}) = \frac{1}{w_{i_h}^2} V(\beta_{i_h})$$

Standard errors for all elasticities are obtained from square root of the variances.

Dynamic AIDS Model

Following Eales and Unnevehr (1988) in considering the importance of dynamics in meat demand, the dynamic AIDS specification of source differentiated U.S. meat demand uses the first difference form of the AIDS model as follows

$$(17) \quad \Delta w_{i_h} = \sum_j \sum_k r_{i_h j_k} \Delta \ln p_{j_k} + b_{i_h} \Delta \ln \frac{E}{P^*}$$

where u is an intercept which represents a trend variable. To avoid simultaneity, like the AIDS model, the lagged expenditure share is used for the Stone's price index.

Homogeneity and Slutsky symmetry are imposed on this system;

$$(17.1) \quad \sum_j \sum_k r_{i_h j_k} = 0$$

$$(17.2) \quad r_{i_h j_k} = r_{j_k i_h}$$

Rotterdam Model

The Rotterdam model is in differential form (Theil, 1965; Barten, 1969). The Rotterdam model starts with the following double-log specification

$$(18) \quad \ln q_{i_h} = a_{i_h} + \sum_j \sum_k e_{i_h j_k} \ln p_{j_k} + \eta_{i_h} \ln E$$

where q_{i_h} denotes the quantity of good i from source h, p_{j_k} denotes the price of good j from source k, η_{i_h} denotes the expenditure elasticity for good i from source h, and $e_{i_h j_k}$ is the cross-price elasticity of good k from source j on the demand of good i from source h. Total differentiation of the double-log specification yields

(19) $$d\ln q_{i_h} = \sum_j \sum_k e_{i_h j_k} \, d\ln_{j_k} + \eta_{i_h} \, d\ln E$$

Substituting the Slutsky decomposition, $e_{i_h} = e^*_{i_h j_k} - \eta_{i_h} w_{j_k}$ where $e^*_{i_h j_k}$ is the compensated cross-price elasticity, into the above equation yields

(20) $$d\ln q_{i_h} = \sum_j \sum_k e^*_{i_h j_k} \, d\ln p_{j_k} + \eta_{i_h} \, (d\ln E - \sum_j \sum_k w_{j_k} dp_{j_k})$$

The above equation does not readily lend itself to the imposition of Slutsky symmetry, $w_{i_h} e^*_{i_h j_k} = w_{j_k} e^*_{j_k i_h}$, since the restrictions also involve variable expenditure shares. This is avoided by multiplying by the expenditure share w_{i_h} , and the Rotterdam model is finally obtained:

(21) $$w_{i_h} d\ln q_{i_h} = \theta_{i_h} \, d\ln Q + \sum_j \sum_k \phi_{i_h j_k} d\ln p_{j_k}$$

where

$$d\ln Q = d\ln E - \sum_j \sum_k w_{j_k} d\ln p_{j_k} = \sum_j \sum_k w_{j_k} d\ln q_{j_k}$$

$$\theta_{i_h} = w_{i_h} \eta_{i_h} = p_{i_h} \frac{\partial q_{i_h}}{\partial E}$$

$$\phi_{i_h j_k} = w_{i_h} e^*_{i_h j_k} = \frac{p_{i_h} p_{j_k} s_{i_h j_k}}{E}$$

where $s_{i_h j_k}$ is the $(i_h j_k)^{th}$ term of the Slutsky substitution matrix, and dlnQ is a Divisia index representing the proportional change in real total expenditure.

For estimation purposes, the parameters θ_{i_h} *and* $\phi_{i_h j_k}$ are treated as constant parameters and the first difference is used to approximate the differential. The Rotterdam model is thus obtained:

$$(22) \quad \bar{w}_{i_h t}\, DX_{i_h t} = \theta_{i_h} DQ_t + \sum_j \sum_k \phi_{i_h j_k} DP_{j_k t}$$

where t denotes time, $\bar{w}_{i_h t} = \frac{(w_{i_h t} + w_{i_h t-1})}{2}$, $DZ_t = \ln(Z_t) - \ln(Z_{t-1})$ where Z represents X, P, and Q, and $DQ_t = \sum_j \sum_k w_{j_k t} d\ln q_{j_k t}$ which is the Divisia quantity index for a given food group. Homogeneity and Symmetry conditions, respectively, require that

$$(23) \quad \sum_j \sum_k \phi_{i_h j_k} = 0 \ ; \quad \phi_{i_h j_k} = \phi_{j_k i_h}$$

Slutsky (compensated) price elasticities for the Rotterdam model are

$$(24) \quad C^*_{i_h j_k} = \frac{\phi^*_{i_h j_k}}{\bar{w}^{**}_{i_h}}$$

where $\bar{w}^{**}_{i_h}$ is the average expenditure share of good i from source h in total expenditure for a given food group. The variance of compensated price elasticities is

$$V(C^*_{i_h j_k}) = \left(\frac{1}{\bar{w}^{**}_{i_h}}\right)^2 V(\phi^*_{i_h j_k})$$

The Cournot (uncompensated) price elasticities are

$$(25) \quad C_{i_h j_k} = \frac{\phi^*_{i_h j_k}}{\bar{w}^{**}_{i_h}} - \frac{\theta^*_{i_h} \bar{w}^{**}_{j_k}}{\bar{w}^{**}_{i_h}}$$

The variance of uncompensated price elasticities is

$$V(C_{i_h j_k}) = \left(\frac{1}{\bar{w}^{**}_{i_h}}\right)^2 V(\phi^*_{i_h j_k}) + \left(\frac{\bar{w}^{**}_{j_k}}{\bar{w}^{**}_{i_h}}\right) V(\theta^*_{i_h}) - 2\left(\frac{\bar{w}^{**}_{j_k}}{\bar{w}^{**2}_{i_h}}\right) Cov(\phi^*_{i_h j_k}, \theta^*_{ih})$$

The expenditure elasticity is

$$(26) \quad \eta^*_{i_h} = \frac{\theta^*_{i_h}}{\overline{W}^{**}_{i_h}}$$

The variance of expenditure elasticities is

$$V(\eta^*_{i_h}) = \left(\frac{1}{\overline{W}^{**}_{i_h}}\right)^2 V(\theta^*_{i_h})$$

Standard errors for all elasticities are obtained from square roots of the variances.

Endogeneity Test

In estimating demand systems such as the AIDS or the Rotterdam model, expenditure, one of explanatory variables, may be endogenous because expenditures are used to compute the dependent variable (Attfield, 1985; Lafrance, 1991). the endogenous expenditure is correlated with error terms and the correlation with the error term causes estimates to be biased and inconsistent. Most previous literature assumes that simultaneity is small and ignores the problem. We follow Blundell (1987) and use the Wu-Hausman test to determine if expenditure can be treated as exogenous.

Let V_{i_h} *and* Z_{i_h} be error terms in the AIDS and the Rotterdam models, respectively. For the purpose of testing the exogeneity assumption of expenditure the equations for $\ln(\frac{E_{i_h}}{P^*})$ in the AIDS and DQ in the Rotterdam model are approximated by

$$(27) \quad \ln \left(\frac{E}{P^*}\right)_t = a_{i_h} + \sum_j \sum_k f_{i_h j_k} \ln P_{j_k t} + g_{i_h} \ln \left(\frac{E}{P^*}\right)_{t-1} + h_{i_h} \ln Y_t$$

$$(28) \quad DQ_t = b_{i_h} + c_{i_h} DQ_{t-1} + \sum_j \sum_k d_{i_h j_k} DP_{j_k t} + DY_t + Z^*_{i_h t}$$

where t is time, Y is total income (GNP is used in this paper), and $V^*_{i_h t}$ and $Z^*_{i_h t}$ are random error terms. The random error terms are partitioned as follows

$$(29) \quad V_{i_h} = \xi V^*_{i_h} + e_{i_h}$$
$$(30) \quad Z_{i_h} = \zeta Z^*_{i_h} + \upsilon_{i_h}$$

where ξ_{i_h} and ζ_{i_h} are correlation parameters such that $E(V^*_{i_h}, e_{i_h}) = 0$ and $E(Z^*_{i_h}, \upsilon_{i_h}) = 0$ and thus e_{i_h} and υ_{i_h} are independent of $V^*_{i_h}$ and $Z^*_{i_h}$, respectively. Then the residuals $V^*_{i_h}$ and $Z^*_{i_h}$ are included in the AIDS equation and the Rotterdam equation, respectively. The F-tests for the inclusion of the residuals provide an asymptotically efficient exogeneity test.

<u>Separability Test</u>

Weak separability applies when the marginal rate of substitution between any two goods in the same group is independent of quantities consumed outside the group. Weak separability is a necessary and sufficient condition for multi-stage budgeting.

For the separability tests, we follow Hayes, Wahl and Williams's test which is based on quasi separability of the cost function. Pudney (1981) argued the different definitions of quasi and weak separability have made little difference to the empirical results. Two groups, i and j, may be considered separable if the compensated cross-price effects between the share of good i from source h in group i and the price of good j from source k in group j ($i \neq j$) satisfy the following restriction

$$(31) \quad \psi_{i_h j_k} = w^*_{i_h} w^*_{j_k} \psi_{ij}$$

where $w^*_{i_h}$ is the intragroup budget share $\frac{p_{i_h} q_{i_h}}{E_i}$ in which E_i is expenditure for group i. The above restriction can be written as

$$(32) \quad \psi_{i_h j_k} \left(\frac{1}{w^*_{i_h} w^*_{j_k}}\right) = \psi_{ij}$$

ψ is γ for the AIDS model and the dynamic AIDS model, and ϕ for the Rotterdam model. To implement the restrictions, Hayes et al. implicitly assume ψ_{ij} is constant. If h is from 1 to 3 and k is from 1 to 2, then six restrictions are required based on equation (32). In this study, however, ψ_{ij} is not assumed constant and five restrictions are used as

$$(33) \quad \frac{\psi_{i_1 j_1}}{w^*_{i_1} w^*_{j_1}} = \frac{\psi_{i_1 j_2}}{w^*_{i_1} w^*_{j_2}} = \frac{\psi_{i_2 j_1}}{w^*_{i_2} w^*_{j_1}} = \frac{\psi_{i_2 j_2}}{w^*_{i_2} w^*_{j_2}} = \frac{\psi_{i_3 j_1}}{w^*_{i_3} w^*_{j_1}} = \frac{\psi_{i_3 j_2}}{w^*_{i_3} w^*_{j_2}}$$

If group j includes one good and thus $w^*_{j_k} = 1$, then restriction (32) is rewritten as

$$(34) \quad \psi_{i_h j} \left(\frac{1}{w^*_{i_h}}\right) = \psi_{ij}$$

Then the case of restriction (33) changes to two restrictions and is rewritten as

$$(35) \quad \frac{\psi_{i_1 j}}{w^*_{i_1}} = \frac{\psi_{i_2 j}}{w^*_{i_2}} = \frac{\psi_{i_3 j}}{w^*_{i_3}}$$

Table 5 shows utility trees for separability tests in this study. In this study, the separability between imported meat vs. domestic meat, source differentiated import beef vs. other meat, source differentiated import beef vs. other import meat, source differentiated import beef vs. domestic beef is tested.

DATA

Annual time series data from 1974 to 1991 are used. Source differentiated U.S. beef import data in value and quantity are obtained from Foreign Agricultural Trade of the United States (FATUS), USDA/ERS. Beef import quantities from FATUS are reported in product weight. The equivalent carcass weights for total beef imports are available from the Livestock and Poultry Situation and Outlook Report (LPSOR), USDA/ERS, but for source differentiated beef imports data is available only after 1983. Thus all such data in product weight prior to 1983 are converted to carcass weights using the following conversion equation

(36) CW = A + B*PW

where CW and PW denote carcass weight and product weight, and the parameters A and B are estimated using the data of carcass and product weight from 1983 to 1991.

Fed beef carcasses in pounds are obtained from the following

(37) Fed Beef = Fed Slaughter * Average Dressed Weight

where Fed Slaughter is fed cattle slaughter in head, and

$$(38)\ AvgDrWT = StDrWT\left(\frac{StSltr}{StSltr+HfSltr}\right) + HfDrWT\left(\frac{HfSltr}{StSltr+HfSltr}\right)$$

where AvgDrWT denotes average dressed weight in pounds, StDrWT is steer dressed weight in pounds, StSlpr is steer slaughter in head, HfDrWT is heifer dressed weight in pounds, HfSltr is heifer slaughter in head. Then the domestic consumption of fed beef is obtained from subtracting total U.S. beef exports in carcass weight from fed beef carcasses. Nonfed beef carcasses in pounds are obtained from the following equation:

(39) Nonfed Beef = (Nonfed Sltr * AvgDrWT) + (Cow Sltr * CowDrWT)

where nonfed Sltr denotes nonfed cattle slaughter in head, Cow Sltr is cow cattle slaughter in head, and CowDrWT is cow dressed weight in pounds. Fed, nonfed, and cow cattle slaughter are obtained from LPSOR, USDA/ERS. Steer and heifer slaughter; steer, heifer, and cow dressed weights are obtained from Livestock Slaughter, USDA/NASS.

Pork consumption in carcass weight is obtained from subtracting total U.S. pork exports in carcass weight from total U.S. commercial pork production in carcass weight. All units are pounds. U.S. commercial pork production, pork exports, and pork imports are obtained from LPSOR, USDA/ERS. Broiler consumption is U.S. ready-to-cook young chicken consumption obtained from LPSOR, USDA/ERS.

For import prices of beef and pork, the unit values obtained by dividing the import value by the carcass weight import quantity are used. For fed and nonfed beef prices, central U.S. wholesale choice beef prices and central U.S. wholesale cow beef prices, respectively, are used. Central U.S. wholesale pork prices are used for domestic pork prices. For broiler prices, a U.S. 12 city composite weighted average wholesale young chicken price is used. The prices for fed and nonfed beef, pork, and broilers are obtained from LPSOR, USDA/ERS.

Gross national Product (GNP), used as total income in the Wu-Hausman test procedure, is obtained from International Financial Statistics, the International Monetary Fund, United Nations.

EMPIRICAL RESULTS

The estimated uncompensated and compensated elasticities of the AIDS, the dynamic AIDS, and the Rotterdam models are reported in Table 4. All system weighted R^2 values are large; 0.98 for the AIDS, 0.96 for the dynamic AIDS, and 0.99 for the Rotterdam model. The compensated own price elasticities of the AIDS model are all significant at the 5-percent significance level except for broilers and ROW beef. The elasticities of the dynamic AIDS model are all significant at the 5-percent significance level except for broilers. The elasticities of the Rotterdam model are also all significant at the 5-percent significance level except for ROW beef. The signs of the elasticities for all three models are reasonable except for Canadian beef.

Tests of theoretical demand conditions of homogeneity and symmetry are reported in Table 3. Homogeneity and symmetry conditions with the Rotterdam model are not rejected, but are rejected for the AIDS model. In the dynamic AIDS model homogeneity is not rejected but symmetry is rejected. However, symmetry imposed on only the three imported beef equations is not rejected in the AIDS and the dynamic AIDS models. In the AIDS model, homogeneity for the three equations is not rejected.

Separability Tests

Table 6 reports separability tests using the AIDS, dynamic AIDS, and Rotterdam models. The separability tests show how the United States allocates its budget on meat among different types of meat. Separability between source differentiated imported beef vs. domestic fed beef and nonfed beef is rejected at the 5-percent level of significance in all models, while separability between source differentiated imported beef vs. domestic pork and broilers is not rejected in all models. Separability between source differentiated imported beef and imported pork is not rejected at the 5-percent level of significance in the dynamic AIDS and Rotterdam models, but is rejected in the AIDS model.

The results of separability tests indicate that the assumption of block independence between domestic and imported beef is not appropriate in U.S. import demand estimation for source differentiated beef. That is, demand for U.S. beef imports can not be estimated independently of demand for domestic beef. However, the tests suggest that U.S. beef import demand can be estimated independently of other imported meat and other domestic meat.

Model Selection

According to the results of the separability tests, source differentiated U.S. beef demand systems consist of domestic fed and nonfed beef, Canadian beef, Oceanian beef, and ROW beef. The separability between beef and all other meat is tested and rejected at the 5-percent significance level (Table 6).

The compensated and uncompensated elasticities of the AIDS, the dynamic AIDS, and the Rotterdam models are estimated and reported in Table 8. System weighted R^2 are 0.69 for the AIDS, 0.77 for the dynamic AIDS, and 0.998 for the Rotterdam model.

Tests of theoretical demand conditions of homogeneity and symmetry are reported in Table 7. Tests of homogeneity and symmetry conditions with the Rotterdam model and the dynamic AIDS model are not rejected, but are rejected for the AIDS model.

In the Wu-Hausman tests, the F values are 0.41 with 4 and 42 degrees of freedom for the AIDS, 1.18 with 4 and 38 degrees of freedom for the dynamic AIDS, and 0.88 with 4 and 42 degrees of freedom for the Rotterdam model. Therefore, the Wu-Hausman tests cannot reject the hypothesis that expenditures in the AIDS, dynamic AIDS, and Rotterdam models are not correlated with the error terms. The test results indicate that U.S. expenditures on beef can be treated as exogenous in the models in this paper.

The Rotterdam model seems to be the best fit for the data of U.S. demand for source differentiated beef, in the context of having the highest system weighted R^2 and holding all theoretical demand properties. Therefore, the Rotterdam model is used for an analysis of U.S. demand for source differentiated beef.

Elasticities of Source Differentiated U.S. Beef Demand

The uncompensated and compensated own price elasticities of the Rotterdam model are all significant at the 5-percent significance level and reasonable in sign except for Canadian beef and ROW beef which are not significant at the 5-percent significance level. The compensated own price elasticity for beef is -0.36 for fed beef, -1.81 for nonfed beef, and -1.18 for Oceanian beef.

The results indicate that nonfed beef is elastic to price changes and thus will be highly affected by price changes, while fed beef will be slightly affected by price changes. Imported beef from Oceania is elastic to price changes and thus the percentage response in quantity of Oceanian beef demanded by the United States will be more than the percentage changes in price. The elastic own price elasticities for nonfed beef and Oceanian beef indicate that all nonfed beef are elastic to price changes in the United States because most imported beef from Oceania also are nonfed beef.

Cross price elasticities show competitive relations among products. Positive cross price elasticities indicate competitive relations, while negative cross price elasticities indicate complementary relations. As indicated by the positive and significant compensated cross price elasticities at the 95-percent confidence level, nonfed beef has a competitive relationship with fed beef, Oceanian beef, and beef from ROW in the U.S. beef market. The negative and significant cross price elasticities at the 95-percent confidence level indicate complementary relationships between fed beef vs. beef from ROW, Canadian beef vs. beef from ROW in the U.S. beef market.

The expenditure elasticities are significant at the 5-percent significance level for fed beef and nonfed beef, and at the 10-percent significance level for ROW beef. The elasticities are 0.58 for fed beef, 2.97 for nonfed beef, and 0.72 for ROW beef. For Canadian beef and Oceanian beef, the expenditures are not significant even at the 10-percent significance level. The estimated expenditure elasticities show that nonfed beef is elastic to the expenditure while fed beef is inelastic. These results indicate that the expenditure share for nonfed beef will increase with increasing U.S. expenditure on beef. Among imported beef, ROW beef is inelastic to the expenditure and the percentage response in quantity demanded will be less than that in expenditure.

SUMMARY AND CONCLUSIONS

Recently, the AIDS, dynamic AIDS, and Rotterdam models have been widely used for source differentiated agricultural trade analysis. All previous studies on source differentiated import demand have assumed separability between domestic and imported goods in which demands for imported goods are estimated conditional on a given total import expenditure and independently of demand for domestic goods. Separability is a strong assumption because a decision for import is often based on changes in domestic demand. Therefore, this paper tested the separability between U.S. domestic beef demand and import beef demand differentiated by source using the AIDS, dynamic AIDS, and Rotterdam models. The separability test results suggest that U.S. beef import demand differentiated by source is not independent of domestic beef demand, while that can be estimated independently of other domestic meat.

Therefore, domestic fed and nonfed beef, and imported beef including Canadian beef, Oceanian beef, and ROW beef were used for modeling U.S. beef demand differentiated by source. As a result of comparing three different demand functional forms in terms of system weighted R^2 and theoretical demand properties, the Rotterdam model fitted the data for U.S. beef demand differentiated by source better than the AIDS and dynamic AIDS models.

The empirical results of own price elasticities indicate that nonfed beef will be highly affected by price changes, while fed beef will be slightly affected by price changes. Therefore, if world beef prices increase due to beef trade liberalization, nonfed beef producers may have a loss in the U.S. beef market, but a benefit to fed beef producers may depend on export demand for U.S. fed beef.

The estimated cross price elasticities indicate that nonfed beef competes with fed beef, ROW beef, and Oceanian beef in the U.S. beef market.

The results from estimated expenditure elasticities indicate that with increasing U.S. total expenditure on beef, nonfed beef will have a relatively higher value share in U.S. beef market. Among source differentiated imported beef, value share for ROW beef will increase with increasing U.S. total expenditure on beef, but the percentage response in quantity will be less than price changes.

In choosing the best functional form for source differentiated U.S. beef, this study was limited to comparing R^2 and testing theoretical demand properties across all alternative functional forms. As further work on selecting the best model, non-nested tests for the AIDS, dynamic AIDS, and Rotterdam models are suggested.

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Figure 1. Utility Tree for Meat Consumption in the United States

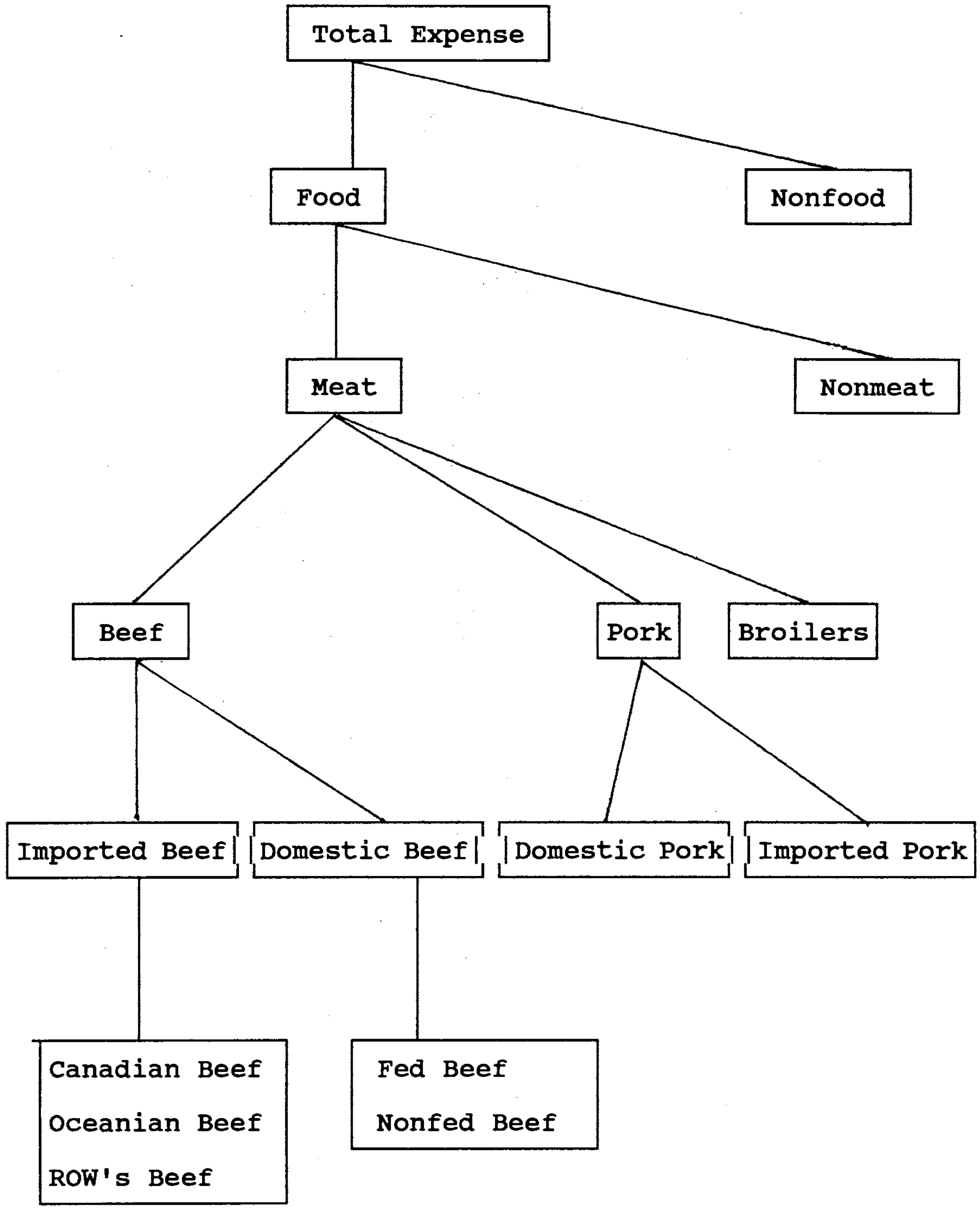

Table 1. Trends in the Value share of World Beef Imports: 1970-1990

Year	Major Importers		
	EC*	U.S.	Japan
1970	0.53	0.31	0.01
1975	0.58	0.16	0.02
1980	0.50	0.18	0.05
1985	0.49	0.15	0.07
1990	0.49	0.13	0.13
Average	0.52	0.19	0.06

* EC denotes the European Community as EC-12 unit throughout.

Source: FAO Trade Yearbook, Food and Agriculture Organization (FAO) of the United Nation, various issues.

Table 2. Summary Statistics for U.S. Meat Expenditure shares: 1974-1991

	Wbc[1]	Wbo[2]	Wbr[3]	Wfb[4]	Wnfb[5]	Wp[6]	Wip[7]	Wbro[8]
Mean	0.0025	0.022	0.009	0.370	0.104	0.331	0.016	0.146
St.Dv.	0.001	0.004	0.002	0.019	0.025	0.014	0.004	0.026
Min	0.0004	0.015	0.007	0.330	0.072	0.308	0.011	0.112
Max	0.004	0.032	0.012	0.401	0.158	0.353	0.024	0.194

[1] Wbc denotes U.S. expenditure share for imported beef from Canada.
[2] Wbo denotes U.S. expenditure share for imported beef from Oceania.
[3] Wbr denotes U.S. expenditure share for imported beef from the rest of the world.
[4] Wfb denotes U.S. expenditure share for domestic fed beef.
[5] Wnfb denotes U.S. expenditure share for domestic nonfed beef.
[6] Wp denotes U.S. expenditure share for domestic pork.
[7] Wip denotes U.S. expenditure share for imported pork.
[8] Wbro denotes U.S. expenditure share for domestic broilers.

Table 3. F-Test Results[1] for Homogeneity and Symmetry

	AIDS	DAIDS[2]	Rotterdam
Homogeneity and symmetry for all equations:			
Homogeneity	8.99** (7,49)	0.38 (7,49)	0.24 (7,56)
Symmetry	5.88** (21,49)	2.18** (21,49)	0.86 (21,56)
Homogeneity for all equations and symmetry for three imported beef equations:			
Homogeneity	8.99** (7,49)	0.38 (7,49)	
Symmetry	0.39 (3,49)	0.41 (3,49)	
Homogeneity and symmetry for three imported beef equations:			
Homogeneity	1.01 (3,49)		
Symmetry	0.39 (3,49)		

[1] Degrees of freedom are in parentheses.
** denotes statistical significance at the 5 percent level.
[2] DAIDS denotes the dynamic AIDS.

Table 4. Compensated and Uncompensated Elasticities[1] from SUR Estimations of Dynamic AIDS and Rotterdam Models

	AIDS		Dynamic AIDS		Rotterdam	
	Marsh	Hicks	Marsh	Hicks	Marsh	Hicks
Fed Beef						
Fed Beef	-0.68	-0.25	-0.88	-0.52	-0.91	-0.56
	(0.10)	(0.09)	(0.29)	(0.14)	(0.17)	(0.08)
Nonfed Beef	-0.08	0.04	0.21	0.31	0.15	0.25
	(0.05)	(0.05)	(0.07)	(0.11)	(0.06)	(0.07)
Domestic Pork	-0.20	0.19	-0.12	0.20	-0.03	0.28
	(0.06)	(0.05)	(0.18)	(0.07)	(0.11)	(0.05)
Imported Pork	0.05	0.07	0.01	0.03	-0.01	0.002
	(0.03)	(0.03)	(0.03)	(0.03)	(0.02)	(0.02)
Broilers	-0.20	-0.03	-0.14	0.001	-0.14	-0.001
	(0.05)	(0.05)	(0.06)	(0.07)	(0.04)	(0.005)
Canadian Beef	-0.005	-0.002	0.005	0.01	0.001	0.003
	(0.004)	(0.004)	(0.01)	(0.01)	(0.004)	(0.004)
Oceanian Beef	-0.01	0.02	-0.02	0.004	0.02	0.04
	(0.03)	(0.03)	(0.02)	(0.04)	(0.03)	(0.03)
ROW Beef	-0.04	-0.03	-0.03	-0.03	-0.04	-0.03
	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)
Expenditure	1.16		0.97		0.95	
	(0.09)		(0.44)		(0.32)	
Nonfed Beef						
Fed Beef	0.83	0.15	-0.13	1.09	0.44	0.90
	(0.24)	(0.19)	(0.92)	(0.39)	(0.65)	(0.26)
Nonfed Beef	-1.07	-0.15	-2.63	-2.29	-2.25	-2.01
	(0.17)	(0.18)	(0.31)	(0.41)	(0.28)	(0.36)
Domestic Pork	1.16	0.55	-0.52	0.56	0.03	0.44
	(0.15)	(0.14)	(0.69)	(0.24)	(0.49)	(0.17)
Imported Pork	-0.15	-0.18	-0.11	-0.06	-0.09	-0.07
	(0.09)	(0.09)	(0.11)	(0.11)	(0.09)	(0.09)
Broilers	0.64	0.37	-0.32	0.15	-0.14	0.03
	(0.12)	(0.11)	(0.23)	(0.22)	(0.19)	(0.02)
Canadian Beef	-0.08	-0.08	-0.05	-0.04	-0.08	-0.07
	(0.02)	(0.02)	(0.03)	(0.03)	(0.02)	(0.02)
Oceanian Beef	0.48	0.44	0.38	0.45	0.28	0.30
	(0.11)	(0.11)	(0.17)	(0.16)	(0.13)	(0.13)
ROW Beef	0.03	0.02	0.11	0.13	0.06	0.07
	(0.02)	(0.02)	(0.04)	(0.03)	(0.03)	(0.02)
Expenditure	-1.85		3.30		2.30	
	(0.25)		(1.61)		(1.47)	

Domestic Pork						
Fed Beef	-0.08	0.21	-0.12	0.23	-0.10	0.32
	(0.07)	(0.06)	(0.18)	(0.07)	(0.11)	(0.05)
Nonfed Beef	0.09	0.17	0.08	0.18	0.02	0.14
	(0.04)	(0.04)	(0.05)	(0.07)	(0.04)	(0.05)
Domestic Pork	-0.75	-0.50	-0.70	-0.39	-0.79	-0.39
	(0.07)	(0.06)	(0.17)	(0.07)	(0.13)	(0.06)
Imported Pork	0.07	0.08	0.01	0.02	-0.0002	0.02
	(0.03)	(0.03)	(0.03)	(0.03)	(0.03)	(0.03)
Broilers	-0.06	0.05	-0.13	0.01	-0.16	0.003
	(0.04)	(0.04)	(0.05)	(0.05)	(0.03)	(0.004)
Canadian Beef	-0.003	-0.001	-0.01	-0.01	-0.01	-0.01
	(0.004)	(0.004)	(0.01)	(0.01)	(0.004)	(0.004)
Oceanian Beef	-0.04	-0.02	-0.05	-0.03	-0.07	-0.04
	(0.02)	(0.02)	(0.02)	(0.02)	(0.03)	(0.03)
ROW Beef	-0.01	0.001	-0.01	-0.004	-0.01	-0.001
	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)
Expenditure	0.78		0.92		1.22	
	(0.08)		(0.34)		(0.27)	
Imported Pork						
Fed Beef	0.34	1.64	-0.73	0.65	-0.64	0.04
	(0.92)	(0.72)	(1.65)	(0.80)	(1.16)	(0.56)
Nonfed Beef	-1.54	-1.17	-0.80	-0.42	-0.68	-0.49
	(0.54)	(0.59)	(0.63)	(0.73)	(0.55)	(0.59)
Domestic Pork	0.53	1.69	-0.72	0.51	-0.24	0.37
	(0.65)	(0.68)	(1.21)	(0.59)	(1.09)	(0.64)
Imported Pork	-1.62	-1.57	-0.31	-0.26	-0.20	-0.17
	(0.48)	(0.47)	(0.50)	(0.51)	(0.52)	(0.53)
Broilers	-1.03	-0.52	-0.96	-0.42	-0.28	-0.02
	(0.45)	(0.44)	(0.48)	(0.48)	(0.31)	(0.04)
Canadian Beef	-0.13	-0.13	-0.01	-0.002	-0.001	0.004
	(0.07)	(0.07)	(0.09)	(0.09)	(0.07)	(0.07)
Oceanian Beef	-0.38	-0.30	-0.35	-0.27	-0.09	-0.05
	(0.34)	(0.34)	(0.46)	(0.47)	(0.44)	(0.45)
ROW Beef	0.31	0.34	0.16	0.20	0.31	0.33
	(0.08)	(0.08)	(0.11)	(0.11)	(0.09)	(0.09)
Expenditure	3.52		3.71		1.91	
	(0.92)		(2.65)		(2.37)	

Broilers

Fed Beef	-1.21	-0.08	0.06	0.002	-0.04	-0.004
	(0.15)	(0.13)	(0.40)	(0.18)	(0.06)	(0.01)
Nonfed Beef	-0.06	0.26	0.12	0.11	0.03	0.04
	(0.08)	(0.08)	(0.11)	(0.16)	(0.02)	(0.02)
Domestic Pork	-0.90	0.12	0.06	0.01	-0.03	-0.0004
	(0.11)	(0.09)	(0.30)	(0.12)	(0.05)	(0.01)
Imported Pork	-0.10	-0.06	-0.04	-0.05	-0.004	-0.002
	(0.05)	(0.05)	(0.05)	(0.05)	(0.004)	(0.004)
Broilers	-0.65	-0.20	-0.06	-0.08	-0.06	-0.04
	(0.12)	(0.12)	(0.12)	(0.14)	(0.025)	(0.004)
Canadian Beef	0.01	0.01	0.01	0.01	-0.001	-0.001
	(0.01)	(0.01)	(0.01)	(0.01)	(0.001)	(0.001)
Oceanian Beef	-0.12	-0.05	-0.01	-0.01	0.01	0.01
	(0.04)	(0.04)	(0.06)	(0.06)	(0.01)	(0.01)
ROW Beef	-0.03	-0.002	0.003	0.002	-0.0004	0.001
	(0.01)	(0.01)	(0.02)	(0.01)	(0.002)	(0.001)
Expenditure	3.06		-0.15		0.15	
	(0.18)		(0.66)		(0.16)	

Canadian Beef

Fed Beef	-2.34	-0.30	0.02	1.07	0.21	0.64
	(0.81)	(0.66)	(2.23)	(1.06)	(1.31)	(0.84)
Nonfed Beef	-4.07	-3.50	-1.87	-1.57	-3.15	-3.09
	(0.66)	(0.70)	(1.13)	(1.24)	(0.78)	(0.78)
Domestic Pork	-1.96	-0.13	-2.29	-1.35	-1.07	-0.87
	(0.54)	(0.57)	(1.61)	(0.76)	(1.04)	(0.59)
Imported Pork	-0.88	-0.79	-0.06	-0.01	0.02	0.03
	(0.42)	(0.41)	(0.56)	(0.56)	(0.46)	(0.46)
Broilers	0.04	0.84	0.32	0.73	-0.14	-0.06
	(0.40)	(0.41)	(0.63)	(0.67)	(0.35)	(0.04)
Canadian Beef	4.77	4.79	4.86	4.87	4.77	4.78
	(0.45)	(0.45)	(0.50)	(0.50)	(0.62)	(0.62)
Oceanian Beef	-0.12	0.005	-1.96	-1.89	-0.44	-0.42
	(0.91)	(0.91)	(1.11)	(1.12)	(0.78)	(0.80)
ROW Beef	-0.96	-0.91	-1.87	-1.85	-2.39	-2.38
	(0.34)	(0.34)	(0.39)	(0.38)	(0.30)	(0.29)
Expenditure	5.52		2.83		4.55	
	(0.74)		(3.54)		(3.10)	

Oceanian Beef

Fed Beef	0.38	0.26	0.83	0.08	0.60	0.66
	(0.52)	(0.44)	(0.97)	(0.36)	(0.98)	(0.45)
Nonfed Beef	2.16	2.12	2.32	2.11	1.41	1.43
	(0.49)	(0.51)	(0.75)	(0.77)	(0.67)	(0.63)
Domestic Pork	-0.18	-0.28	0.28	-0.39	-0.72	-0.66
	(0.36)	(0.36)	(0.88)	(0.35)	(0.72)	(0.41)
Imported Pork	-0.21	-0.21	-0.16	-0.19	-0.04	-0.03
	(0.25)	(0.24)	(0.33)	(0.34)	(0.31)	(0.32)
Broilers	-0.32	-0.37	0.22	-0.17	-0.03	-0.001
	(0.26)	(0.27)	(0.39)	(0.37)	(0.25)	(0.03)
Canadian Beef	0.001	0.001	-0.20	-0.21	-0.05	-0.05
	(0.10)	(0.10)	(0.12)	(0.12)	(0.09)	(0.09)
Oceanian Beef	-1.89	-1.89	-1.44	-1.49	-1.30	-1.36
	(0.58)	(0.57)	(0.67)	(0.68)	(0.75)	(0.77)
ROW Beef	0.37	0.37	0.17	0.15	0.23	0.23
	(0.12)	(0.12)	(0.14)	(0.14)	(0.13)	(0.13)
Expenditure	-0.31		-2.01		-2.63	
	(0.47)		(1.81)		(1.97)	

ROW Beef

Fed Beef	-1.35	-1.52	-0.97	-1.13	-1.07	-1.17
	(0.28)	(0.23)	(0.67)	(0.33)	(0.47)	(0.23)
Nonfed Beef	0.13	0.08	1.37	1.33	0.89	0.86
	(0.24)	(0.25)	(0.35)	(0.37)	(0.29)	(0.28)
Domestic Pork	0.18	0.03	0.12	-0.03	0.04	-0.05
	(0.21)	(0.22)	(0.46)	(0.22)	(0.36)	(0.21)
Imported Pork	0.40	0.40	0.22	0.22	0.59	0.58
	(0.14)	(0.14)	(0.18)	(0.18)	(0.16)	(0.16)
Broilers	0.23	0.16	0.18	0.11	0.05	0.01
	(0.15)	(0.15)	(0.19)	(0.20)	(0.12)	(0.01)
Canadian Beef	0.20	0.20	-0.26	-0.26	-0.66	-0.66
	(0.15)	(0.15)	(0.15)	(0.15)	(0.08)	(0.08)
Oceanian Beef	0.73	0.72	0.33	0.32	0.59	0.58
	(0.30)	(0.30)	(0.34)	(0.34)	(0.31)	(0.32)
ROW Beef	-0.07	-0.08	-0.55	-0.56	-0.07	-0.07
	(0.14)	(0.14)	(0.15)	(0.15)	(0.11)	(0.11)
Expenditure	-0.46		-0.45		-0.12	
	(0.26)		(1.06)		(0.85)	
System weighted R^2	0.98		0.96		0.99	

[1] Standard errors are in parentheses. Homogeneity and symmetry were imposed for the AIDS, dynamic AIDS, and Rotterdam models. To avoid simultaneity problems, the Stone's price index with lagged shares is used for the AIDS and dynamic AIDS models. For the Rotterdam model, Divisia volume index with lagged shares is used.

Table 5. Possible Utility Trees for Meat Consumption in United States for Separability Test

Commodity	Utility Tree[1] 1	2
Beef		
Domestic Beef		
Fed Beef	A	A
Nonfed Beef	A	A
Imported Beef		
Canadian Beef	B	B
Oceanian Beef	B	B
ROW's Beef	B	B
Pork		
Domestic Pork	A	C
Imported Pork	B	D
Broilers	A	E
No. of Commodity Groups	2	5

[1] In each tree, all commodities with the same letter are assumed to belong to the same group, while commodities with different letters are assumed to be weakly separable.

Table 6. F-Test Results[1] of Separability

Separability	AIDS	Dynamic AIDS	Rotterdam
Imported Meat[2] Separable from			
Domestic Meat[3]	4.28** (15,77)	3.24** (15,84)	1.60* (15,84)
Imported Beef[4] Separable from			
Domestic Beef[5]	4.40** (5,77)	3.53** (5,84)	2.64** (5,84)
Other Domestic Meat[6]	0.57 (5,77)	1.98 (5,84)	0.70 (5,84)
Pork	0.35 (2,77)	1.76 (2,84)	1.19 (2,84)
Broilers	1.06 (2,77)	0.91 (2,84)	0.23 (2,84)
Imported Other Meat[7]	15.14** (2,77)	2.95* (2,84)	1.36 (2,84)

[1] Homogeneity and symmetry are imposed on all equations. Degrees of freedom are in parentheses. ** denotes statistical significance at the 5 percent level, and * at the 10 percent level.

[2] Imported meat includes source differentiated imported beef and imported pork.

[3] Domestic meat includes domestic fed and nonfed beef, domestic pork, and domestic broilers.

[4] Imported beef includes source differentiated imported beef.

[5] Domestic beef includes domestic fed and nonfed beef.

[6] Other domestic meat includes domestic pork and broilers.

[7] Imported other meat includes imported pork.

Table 7. F-Test Results[1] for Homogeneity and Symmetry: U.S. Beef Demand System (1974-1991)

	AIDS	Dynamic AIDS	Rotterdam
Homogeneity	14.87** (4,40)	1.71 (4,36)	0.95 (4,44)
Symmetry	11.59** (6,40)	1.65 (6,36)	1.13 (6,44)

[1] Degrees of freedom are in parentheses.
** denotes statistical significance at the 5 percent level.

Table 8. Compensated and Uncompensated Elasticities[1] from SUR Estimations of U.S. Beef Demand System:1974-91

	AIDS		Dynamic AIDS		Rotterdam	
	Marsh	Hicks	Marsh	Hicks	Marsh	Hicks
Fed Beef						
Fed Beef	-0.70	-0.42	-0.75	-0.25	-0.78	-0.36
	(0.08)	(0.04)	(0.10)	(0.04)	(0.17)	(0.06)
Nonfed Beef	0.39	0.47	0.16	0.30	0.28	0.40
	(0.04)	(0.04)	(0.04)	(0.04)	(0.06)	(0.06)
Canadian Beef	-0.01	-0.01	-0.005	-0.001	-0.005	-0.002
	(0.005)	(0.005)	(0.003)	(0.003)	(0.003)	(0.003)
Oceanian Beef	-0.04	-0.02	-0.06	-0.03	-0.045	-0.02
	(0.01)	(0.01)	(0.02)	(0.02)	(0.016)	(0.01)
ROW Beef	-0.02	-0.01	-0.03	-0.02	-0.03	-0.02
	(0.004)	(0.003)	(0.004)	(0.004)	(0.005)	(0.003)
Expenditure	0.37		0.68		0.58	
	(0.10)		(0.10)		(0.19)	
Nonfed Beef						
Fed Beef	-1.13	1.67	-0.90	1.07	-0.73	1.43
	(0.28)	(0.14)	(0.37)	(0.16)	(0.59)	(0.21)
Nonfed Beef	-2.91	-2.13	-2.00	-1.45	-2.42	-1.81
	(0.18)	(0.17)	(0.19)	(0.21)	(0.23)	(0.24)
Canadian Beef	-0.04	-0.02	-0.08	-0.06	-0.04	-0.02
	(0.04)	(0.04)	(0.02)	(0.02)	(0.03)	(0.03)
Oceanian Beef	0.26	0.43	0.21	0.33	0.17	0.29
	(0.09)	(0.09)	(0.13)	(0.13)	(0.12)	(0.11)
ROW Beef	-0.02	0.05	0.06	0.11	0.06	0.11
	(0.03)	(0.03)	(0.03)	(0.03)	(0.03)	(0.03)
Expenditure	3.83		2.70		2.97	
	(0.34)		(0.37)		(0.67)	
Canadian Beef						
Fed Beef	0.08	-1.63	-0.62	-0.17	-0.97	-0.32
	(1.47)	(0.67)	(1.00)	(0.42)	(1.15)	(0.42)
Nonfed Beef	-0.24	-0.72	-2.73	-2.61	-1.14	-0.96
	(1.52)	(1.46)	(1.02)	(1.03)	(1.05)	(1.07)
Canadian Beef	0.80	0.79	1.69	1.70	1.03	1.03
	(1.46)	(1.46)	(0.72)	(0.72)	(0.81)	(0.81)
Oceanian Beef	-0.56	-0.66	1.39	1.42	1.09	1.13
	(1.16)	(1.13)	(0.80)	(0.79)	(0.76)	(0.74)
ROW Beef	2.25	2.21	-0.35	-0.34	-0.90	-0.88
	(0.79)	(0.78)	(0.37)	(0.37)	(0.43)	(0.43)
Expenditure	-2.34		0.61		0.89	
	(1.71)		(1.02)		(1.26)	

Oceanian Beef						
Fed Beef	0.71	-0.41	0.56	-0.49	0.44	-0.34
	(0.38)	(0.19)	(0.73)	(0.31)	(0.67)	(0.24)
Nonfed Beef	2.36	2.05	1.87	1.58	1.62	1.40
	(0.39)	(0.41)	(0.58)	(0.61)	(0.49)	(0.53)
Canadian Beef	-0.07	-0.08	0.17	0.16	0.13	0.13
	(0.13)	(0.13)	(0.09)	(0.09)	(0.08)	(0.09)
Oceanian Beef	-1.27	-1.33	-1.18	-1.24	-1.13	-1.18
	(0.41)	(0.40)	(0.58)	(0.57)	(0.44)	(0.43)
ROW Beef	-0.20	-0.22	0.02	-0.01	0.01	-0.01
	(0.11)	(0.11)	(0.12)	(0.12)	(0.10)	(0.10)
Expenditure	-1.53		-1.44		-1.06	
	(0.48)		(0.74)		(0.73)	
ROW Beef						
Fed Beef	-1.26	-0.44	-1.04	-0.77	-1.25	-0.73
	(0.29)	(0.14)	(0.39)	(0.17)	(0.39)	(0.14)
Nonfed Beef	0.33	0.56	1.21	1.28	1.17	1.32
	(0.31)	(0.31)	(0.31)	(0.33)	(0.30)	(0.31)
Canadian Beef	0.64	0.65	-0.10	-0.10	-0.26	-0.26
	(0.23)	(0.23)	(0.11)	(0.11)	(0.13)	(0.13)
Oceanian Beef	-0.62	-0.57	-0.04	-0.02	-0.05	-0.02
	(0.29)	(0.28)	(0.30)	(0.30)	(0.27)	(0.26)
ROW Beef	-0.22	-0.20	-0.39	-0.39	-0.32	-0.31
	(0.22)	(0.21)	(0.19)	(0.19)	(0.20)	(0.20)
Expenditure	1.12		0.37		0.72	
	(0.35)		(0.39)		(0.42)	
System weighted R^2	0.69		0.77		0.998	

[1] Standard errors are in parentheses. Homogeneity and symmetry were imposed for all models. To avoid simultaneity problems, the Stone's price index with lagged shares is used for the AIDS and dynamic AIDS models, and the Divisia volume index with lagged shares is used for the Rotterdam model.

VITA

Jung-Hee Lee

Candidate for the Degree of

Doctor of Philosophy

Thesis: NON-NESTED TESTS AND AGRICULTURAL TRADE MODELS

Biographical:

Personal Data: Born in Korea, June 12, 1961, the son of Mr. Neung-Uk Lee and Mrs. Ok-Sun Lee(Shin). Married to Ho-Jin Lee(Chung) on April 26, 1986.

Education: Graduated from Han-Sung High School, Seoul, Korea in February, 1978; received Bachelor of Economics degree from Chung-Ang University, Seoul, Korea in February, 1985; received Master of Science degree in Agricultural Economics at Oklahoma State University in December, 1989; completed requirements for the Doctor of Philosophy Degree at Oklahoma State University in May, 1993.

Professional Experience: Graduate Research Assistant, Department of Agricultural Economics, Oklahoma State University, December 1988 to May 1993. Staff, Division of Food Marketing Management, Lotte Confectionery Co., Ltd. Seoul, Korea, January 1985 to March 1987.

Honors: Recognition from The Honor Society of Agriculture, Gamma Sigma Delta. Graduate Research Excellence Award, Oklahoma State University in May, 1993.