Demand Analysis of the U.S. Fresh Tomato Market

Jione Jung
Postdoctoral Research Associate
E-Mail: jjung@ifas.ufl.edu

John VanSickle
Professor and Director

James Seale
Professor

International Agricultural Trade and Policy Center
Food and Resource Economics
University of Florida
P.O. Box 110240
Gainesville, FL 32611

Selected Paper Prepared for Presentation at the Southern Agricultural Economics Association Annual Meeting, Little Rock, Arkansas, February 5-9, 2005

Demand Analysis of the U.S. Fresh Tomato Market

Abstract

The U.S. fresh tomato industry has been growing significantly over the past several decades. However, as a net importer of fresh tomatoes, the United States imported 36% of total fresh tomato consumption in 2002. The objective of this study is to estimate U.S. demand for domestic and import fresh tomatoes using empirical demand models. Conditional price and expenditure elasticities for U.S. fresh tomato demand are estimated.

Key words: U.S. fresh tomato import, demand system, price and expenditure elasticities

DEMAND ANALYSIS OF THE U.S. FRESH TOMATO MARKET

The U.S. fresh tomato industry has been growing significantly over the past several decades and the United States is currently one of the world's leading fresh tomato producers, ranking second to China. Imports of fresh tomatoes also have risen sharply since 1994. As a net importer of fresh tomatoes, the United States imported 36% of total fresh tomato consumption in 2002. The objective of this study is to estimate U.S. demand for fresh tomatoes using the Rotterdam model and the first-difference version of the almost ideal demand system (AIDS). This analysis examines import demand as well as demand for domestically produced fresh tomatoes. Price and expenditure elasticities for tomatoes imported from Mexico and Canada are estimated to compare those for domestic tomatoes.

Overview of U.S. Fresh Tomato Industry

The increase in tomato production is due to improved efficiency at the grower and processor levels. In 2002, U.S. fresh tomato production totaled 37 million hundred weights (U.S. Department of Agriculture (USDA), Economic Research Service (ERS), July 2003). California and Florida account for two-thirds of the acres used to grow fresh tomatoes in the United States (31% and 39% in 2002, respectively; USDA, ERS, 2003). Tomatoes lead in cash receipts along with potatoes and lettuce, with $1.6 billion in 2001, 10% of all vegetable and melon cash receipts (USDA, ERS, July 2003)

Fresh tomato consumption has continuously increased. During the recent 3 years (2002-02), average fresh tomato consumption was 17.8 pounds per person annually (USDA, ERS, July 2003). The increase in consumption is likely the result of increased consumers' awareness of health and nutrition benefits from eating more tomatoes. One medium fresh tomato (about 5.2 oz) has 35 calories and provides 40% of the U.S. Recommended Daily Allowances (USRDA) of vitamin C and 20% of vitamin A (USDA, ERS, 2000). As a good source of lycopene, fresh tomatoes (and tomato products) may help prevent cancer and heart disease (Florida Tomato Committee).

As a net importer of fresh tomatoes, the United States imported 36% of total U.S. fresh tomato consumption in 2002, while exports accounted for 9% of total domestic production (USDA, ERS, July 2003). In 2002, the United States imported 1.9 billion pounds of fresh tomatoes, valued at $795 million (USDA, Foreign Agricultural Service (FAS)). Mexico, Canada, and the Netherlands are the major fresh tomato exporters to the United States (Figure 1). Mexico and Canada are the dominant fresh tomato exporters to the United States, accounting for 91% of the total value of U.S. fresh tomato imports and 95% of the quantity (USDA, ERS, July 2003; USDA, FAS).

Data

Data on U.S. fresh tomato imports from Mexico, Canada were obtained from the USDA, FAS. Monthly import data from January 1990 to December 2001 are used in this analysis and the sample size contains 144 observations. The quantity of imports from each country is measured in metric tons, and the value of imports is defined as cost insurance freight (c.i.f.) prices. Unit prices of imported fresh tomatoes from each country are derived by dividing total value by total quantity of imports.

Domestic data were obtained from U.S. Tomato Statistics published by USDA, ERS (March 2003). Monthly shipment and grower prices were used for the period from January 1990 to December 2001. A summary of descriptive statistics is presented in Table 1.

Estimation of Empirical Demand Models

Estimation procedures of this study followed the assumption that domestic products are not separable from the imported good (Winters). Accordingly, a demand model includes information of both domestic and imported products. Three different demand models were estimated: double-log demand model, Rotterdam model, and Almost Ideal Demand System (AIDS). Specially, the double-log demand model was used with parameter restrictions in order to check whether the Armington assumptions are appropriate for the fresh tomato data. Theoretical demand restrictions are maintained when each demand model is estimated using seemingly unrelated regressions (SUR) techniques (Kmenta; Zellner).[1]

Double-log demand model

Starting with the following logarithmic demand function,

(1) $$\log q_i = \alpha_i + \sum_j r_{ij} \log p_j + b_i \log E$$

where q_i is the quantity from source i, p_j is the price of products from source j, and E is total expenditure. For the purpose of empirical estimation of elasticities, this model has been frequently applied since parameters of the model themselves present elasticity. The total expenditure elasticities (e_i) and uncompensated price elasticities (e_{ij}) are as the following:

(2) $$e_i = \frac{\partial \log g_i}{\partial \log E} = b_i$$

$$e_{ij} = \frac{\partial \log g_i}{\partial \log p_j} = r_{ij}$$

[1] The command in the TSP program for the first technique is "SUR", while for the latter maximum likelihood technique "LSQ" command is appropriate. Under the assumption of the normality of disturbances, this study applied "LSQ" command for the estimation.

That is, Equation (1) can be rewritten as

(3) $$\log q_i = \alpha_i + \sum_j e_{ij} \log p_j + e_i \log E$$

From the Slutsky equation, the following relationship is true:

(4) $$e_{ij} = e_{ij}^{*} - e_i w_j$$

where e_{ij}^{*} is the compensated cross-price elasticity and w_j is the budget share. Substituting Equation (4) into Equation (3),

(5) $$\log q_i = \alpha_i + \sum_j e_{ij}^{*} \log p_j + e_i \left(\log E - \sum_j w_j \log p_j \right)$$

Let $\sum_j w_j \log p_j$ be $\log P^{*}$ as a price index, then the demand model in Equation (1) can be expressed in terms of real expenditure and compensated prices:

(6) $$\log q_i = \alpha_i + \sum_j e_{ij}^{*} \log p_j + e_i \log\left(\frac{E}{P^{*}} \right)$$

Homogeneity of Equation (3) implies the following:

(7) $$\sum_j e_{ij}^{*} = 0$$

The corresponding elasticities of the model in Equation (6) are

(8) Total expenditure elasticities: $\eta_i = e_i$

Uncompensated cross-price elasticities: $\varepsilon_{ij} = e_{ij}^{*} - e_i w_j$

Compensated cross-price elasticities: $\varepsilon_{ij}^{*} = e_{ij}^{*}$

Now, theoretical demand restrictions are considered particularly for this model. The model itself in Equation (6) is homogeneous of degree zero in all prices and total expenditure (i.e., no need to test homogeneity). In general, the adding-up and symmetry restrictions are not possible to impose into a double-log specification. However, related to the adding-up and symmetry restrictions, an original Armington model is nested within the double-log specification in Equation (6) under the following conditions (Alston et al., 1990):[2]

(9) Demands are homothetic: $e_i = 1 \quad \forall\ i$

Weak separability –
Only the own-price is included: $e_{ij}{}^{*} = 0 \quad \forall\ i \neq j$

Single CES: $e_{ii}{}^{*} = e_{jj}{}^{*} = -\sigma \quad \forall \quad i, j$

where σ is the elasticity of substitution for the system.

The parameter restrictions in (9) were tested to check whether the Armington assumptions are valid for the fresh tomato data using the likelihood ratio (LR) procedures. The LR test is based on the idea that if the restrictions are true, the value of the likelihood function maximized with the restrictions imposed cannot differ too much from the value of the likelihood function maximized without the imposition of the restrictions. Asymptotically, the LR test obtains the following test statistics:

(10) $$\lambda = 2(L_{UR} - L_R) \sim \chi^2{}_{(m)}$$

where L_{UR} and L_R are, respectively, the maximum value of the unrestricted log-likelihood function and the maximum value of the restricted log-likelihood function. The test statistics λ follows the chi-square (χ^2) distribution with the degree of freedom (m) equal to the number of restrictions, which can be determined by subtracting the number of the restricted coefficient from the number of the unrestricted coefficient.

---

[2] An original Armington model is specified as follows: $w_i = b_i^{\sigma}\left(\frac{p_i}{P}\right)^{1-\sigma}$, where P is the price index (Armington).

The result of the LR tests is summarized in Table 2. The Armington restrictions were comprehensively rejected with the chi-square tests. The full Armington restriction was also rejected.[3] When the full Armington restrictions were imposed into the double-log demand model, elasticity of substitution could be estimated. The empirical estimate of elasticity of substitution in this research is 1.1076.

Table 3 presents the SUR estimates of the double-log demand model for fresh tomatoes using monthly data from 1990 to 2001. Equations estimated were demand for U.S. (US), Mexican (MX), and Canadian (CD) tomatoes. Own-price parameters exhibit a statistically significant effect in the Mexican and Canadian tomato demand equations. All six cross price parameters are statistically significant implying that the expenditure shares of all three tomatoes depend on the prices of other commodities. The expenditure variable is statistically significant in all three equations.

Uncompensated (Marshallian) and compensated (Hicksian) price and expenditure elasticities and their variances were calculated from the parameter estimates, and the results are summarized in Tables 4 and 5. Uncompensated elasticities contain both price and income effects, while compensated elasticities only include price effects. Especially, compensated price elasticities hold real income and all other prices constant, and therefore it reflects pure substitution effects (Weatherspoon and Seale). In the double-log demand model in Equation (6), the price coefficients can be interpreted as compensated price elasticities.[4] The elasticities were calculated at the sample mean of each commodity expenditure share. The uncompensated own price elasticities of each tomato show

[3] This research tested whether the Armington assumptions are rejected in a double-log functional form. The authors agree that rejection could be based on the functional form of the double-log demand model but not really based on the data itself.
[4] Only if the price coefficients were restricted to sum to zero, they could be interpreted as compensated elasticities (Alston et al., 2002).

negative signs and are statistically significant. Expenditure elasticities for all three tomatoes are positive and statistically significant at the 5% level indicating that they are normal goods. The compensated own price elasticities for all three tomatoes are also negative. Negative cross-price elasticities indicate that commodities are complements, while positive cross-price elasticities indicate a substitute relationship.

Rotterdam model

The finite-change version of the Rotterdam model (Theil) takes the form (without time subscripts for convenience) as

(11) $$w_i \Delta \log q_i = \theta_i DQ + \sum_j \pi_{ij} \Delta \log p_j \,, \qquad i = 1,2,\ldots,n$$

where w_i represents the value or budget share of commodity i. p_i and q_i are the price and quantity of good i, respectively. DQ is the finite-change version of Divisia volume index:

(12) $$DQ = \sum_i w_i \Delta \log q_i$$

Coefficients θ_i and π_{ij} are given by

(13) $$\theta_i = p_i \frac{\partial q_i}{\partial E},$$

$$\pi_{ij} = s_{ij} \frac{p_i p_j}{E},$$

$$s_{ij} = \frac{\partial q_i}{\partial p_j} + q_j \frac{\partial q_i}{\partial E}$$

where E is total expenditure and $E = E_1 + E_2 + \cdots + E_n$. s_{ij} is the (i, j)th element of the Slutsky substitution matrix. θ_i is the marginal budget share for good i. π_{ij} is known as the Slutsky coefficients of the Rotterdam model. Theoretical demand restrictions can be applied to the parameters of the Rotterdam model:

(14) Adding-up: $\sum_i \theta_i = 1, \qquad \sum_i \pi_{ij} = 0$

Homogeneity: $\sum_j \pi_{ij} = 0$

Symmetry: $\pi_{ij} = \pi_{ji}$

Provided the data add up, the adding-up restriction on the Rotterdam model is automatically satisfied so that the sum of the dependent variables in Equation (11) will equal the first independent variables. The homogeneity and symmetry restrictions can be imposed and tested equation by equation.

The Rotterdam model in Equation (11) was estimated for the fresh tomato monthly data from 1990 to 2001. The homogeneity and symmetry restrictions were imposed. The result of the SUR estimation is summarized in Table 6. Due to symmetry, the bottom half is a mirror image of the top half. The Rotterdam demand system was estimated three times and one equation was dropped for avoiding singularity problem at each time of estimation. Own price variable exhibits a statistically significant effect in demand equations of U.S., Mexican, and Canadian tomatoes. The expenditure variable is statistically significant in all three equations.

Elasticities at the sample mean were calculated from the parameter estimates and the results are in Table 7 and Table 8.[5] Uncompensated own-price elasticities are negative except in CD equation. All-own price parameters along the diagonal in all three equations are negative as expected, implying as the price of fresh tomatoes increases, the amount of

---

[5] Calculated elasticities for the Rotterdam model are conditional since this study used a conditional Rotterdam model.

fresh tomato quantity demanded declines. All compensated own-price elasticities are negative and inelastic (-0.1530, -0.3323, and -0.6507). These results indicate that if the price of U.S. fresh tomatoes drops by 1%, the quantity demanded would increase by 0.15%; if the price of Mexican tomatoes decreases by 1%, the quantity demanded would increase by 0.33%; and if the price of Canadian tomatoes declines by 1%, the quantity demand will rise by 0.65%. Positive compensated cross-price elasticities indicate that U.S and Mexican tomatoes are pairwise substitutes and so are U.S. and Canadian tomatoes. Substitution relationship also exists between Mexican and Canadian tomatoes.

Almost ideal demand system (AIDS)

Deaton and Muellbauer developed the AIDS derived from a specific class of preferences (known as price-independent generalized logarithm (PIGLOG) class):

(15) $$w_i = \alpha_i + \sum_j \gamma_{ij} \log p_j + \beta_i \log\left(\frac{E}{P}\right)$$

where P is a price index derived from the AIDS cost function and defined by

(16) $$\log P = \alpha_0 + \sum_k \alpha_k \log p_k + \frac{1}{2}\sum_k \sum_l \gamma_{kl} \log p_k \log p_l$$

Theoretical demand restrictions require that

(17) Adding-up: $\sum_i \alpha_i = 1$, $\sum_i \beta_i = 0$, $\sum_i \gamma_{ij} = 0$

Homogeneity: $\sum_j \gamma_{ij} = 0$

Symmetry: $\gamma_{ij} = \gamma_{ji}$

The differential form of the AIDS is

(18) $$\Delta w_i = \beta_i \Delta \log\left(\frac{E}{P}\right) + \sum_j \gamma_{ij} \Delta \log p_j$$

To estimate Equation (18) practically, $\Delta \log P$ is replaced by Divisia price index, $\sum_i w_i \Delta \log p_i$:

$$\text{(19)} \qquad \Delta w_i = \beta_i \left(\Delta \log E - \sum_i w_i \Delta \log p_i \right) + \sum_j \gamma_{ij} \Delta \log p_j$$

$$= \beta_i DQ + \sum_j \gamma_{ij} \Delta \log p_j$$

The first-difference version of the AIDS (FD/AIDS) was estimated for fresh tomatoes using monthly data form 1990 to 2001. The FD/AIDS was estimated three times to recover standard errors of the parameters of dropped equation. Table 9 presents the parameter estimates of the restricted FD/AIDS with homogeneity and symmetry restrictions. Calculated uncompensated and compensated elasticities from the parameter estimates of the restricted FD/AIDS are shown in Tables 10 and 11. All own-price elasticities are negative as expected and statistically significant at the 5% significance level. Most of expenditure elasticities are positive except for the MX equation.

U.S. and Mexican fresh tomatoes and U.S. and Canadian fresh tomatoes show pairwise substitute relationship in the restricted FD/AIDS. The compensated cross-price elasticity of Mexican fresh tomatoes with respect to U.S. domestic fresh tomatoes (0.3816) is greater than that of U.S. domestic fresh tomatoes with respect to Mexican fresh tomato imports (0.1794). This indicates that the price of U.S. domestic fresh tomatoes affects the expenditure share of Mexican fresh tomato imports, while the price of Mexican tomato imports does not have such an influence on the expenditure share of U.S. domestic fresh tomatoes.

Summary

In this paper, three empirical demand models were estimated with or theoretical demand restrictions imposed: the double-log demand model, Rotterdam model, and first difference version of the AIDS. Estimation of those demand models allowed calculating price elasticities. Uncompensated and compensated elasticities were calculated at the sample mean using the parameter estimates of each demand model. Which demand model can explain best for the current U.S. fresh tomato industry was not determined. Model choice and specification for the given data are potential tasks to do in advising more efficient policy implications.

Table 1. Descriptive statistics on U.S. consumption of fresh tomatoes by source, January 1990-December 2001

	U.S.	Mexico	Canada
Price ($/cwt)			
Mean	$32	$33	$81
Standard Deviation	15	13	29
Minimum	15	12	28
Maximum	116	83	190
Import Value (1,000$)			
Mean	$64,402	$33,544	$4,803
Standard Deviation	24,440	29,053	6,745
Minimum	16,202	3,123	-
Maximum	150,748	151,399	25,901
Import Quantity (1,000cwt)			
Mean	2,117.54	955.09	67.93
Standard Deviation	643.07	667.43	100.24
Minimum	166.00	161.21	-
Maximum	3,529.00	2,852.24	404.22
Budget share			
Mean	0.6347	0.2919	0.0421
Standard Deviation	0.2007	0.1966	0.0574
Minimum	0.1173	0.0253	0.0000
Maximum	0.9699	0.8769	0.2384

Table 2. Chi-square (χ^2) statistics for hypothesis tests using double-log demand model

	Test Statistics	d.f. [c]
Separability	98.905 *	56
Homotheticity	361.150 *	8
H&S [a]	235.421 *	71
Armington [b]	301.942 *	64

[a] H&S denotes the joint restriction of homothetic separability.

[b] "Armington" denotes the full set of Armington restrictions – homotheticity, separability, and equality of own price coefficients ($\gamma_{ii} = \gamma_{jj} \ \forall \ i, j$).

[c] d.f. (i.e., degree of freedom) for the LR test that equals the number of restrictions are as follows (Alston et al., 1990):

Separability:	$n^2 - n$
Homotheticity:	n
H&S:	n^2
Armington:	$n^2 + n - 1$

* indicates a rejection of the null hypothesis that the model is correct at the 5% significance level ($p = 0.05$).

Table 3. Parameter estimates of double-log demand model

Eq.[a]	Intercept	US	MX	CD	Exp[6]
US	-1.2995	-0.1136	0.3943*	-0.2806*	0.7702*
	(1.8549)	(0.0980)	(0.1013)	(0.1007)	(0.1609)
MX	-8.5920*	-0.3035*	-0.6599*	0.9634*	1.3152*
	(2.8757)	(0.1519)	(0.1571)	(0.1561)	(0.2494)
CD	-76.3482*	1.4585*	1.6064*	-3.0649*	6.8758*
	(6.5589)	(0.3464)	(0.3582)	(0.3560)	(0.5688)

[a] Abbreviations:

Eq.: Equation name
US: Coefficient of U.S. prices
MX: Coefficient of Mexican prices
CD: Coefficient of Canadian prices
Exp: Coefficient of expenditure variable

* indicates a coefficient that is significantly different from zero at the 5% significance level ($p = 0.05$). Estimates in parentheses are standard errors.

Table 4. Uncompensated price and expenditure elasticities for double-log demand model (at the sample mean)

Eq.[a]	US	MX	CD	Exp
US	-0.6140*	0.1590	-0.3152*	0.7702*
	(0.1160)	(0.1197)	(0.0072)	(0.1609)
MX	-1.1580*	-1.0617*	0.9044*	1.3152*
	(0.1798)	(0.1856)	(0.0112)	(0.2494)
CD	-3.0085*	-0.4939	-3.3734*	6.8758*
	(0.4100)	(0.4234)	(0.0255)	(0.5688)

[a] Abbreviations are the same with the previous table.
* indicates a coefficient that is significantly different from zero at the 5% significance level ($p = 0.05$). Estimates in parentheses are standard errors.

Table 5. Compensated price elasticities for double-log demand model (at the sample mean)

Eq.[a]	US	MX	CD
US	-0.1136	0.3943*	-0.2806*
	(0.0980)	(0.1013)	(0.1007)
MX	-0.3035*	-0.6599*	0.9634*
	(0.1519)	(0.1571)	(0.1561)
CD	1.4585*	1.6064*	-3.0649*
	(0.3464)	(0.3582)	(0.3560)

[a] Abbreviations are the same with the previous table.
* indicates a coefficient that is significantly different from zero at the 5% significance level ($p = 0.05$). Estimates in parentheses are standard errors.

[6] As noted by LaFrance (1991), when the expenditure variable is constructed from the price and quantity data (as done in this research), it is correlated with the error term in a quantity- or budget share-dependent demand equation. This means that it may not be appropriate to treat the expenditure variable as exogenous. However, this research treats the expenditure variable as exogenous.

Table 6. Parameter estimates of restricted Rotterdam model with homogeneity and symmetry

Eq.[a]	US	MX	CD	Exp
US	-0.0997*	0.0855*	0.0143*	0.6939*
	(0.0231)	(0.0241)	(0.0063)	(0.0664)
MX		-0.1006*	0.0151*	0.2630*
		(0.0264)	(0.0076)	(0.0707)
CD			-0.0294*	0.0431*
			(0.0082)	(0.0161)

[a] Abbreviations are the same with the previous table.
* indicates a coefficient that is significantly different from zero at the 5% significance level ($p = 0.05$). Estimates in parentheses are standard errors.

Table 7. Uncompensated price and expenditure elasticities for restricted Rotterdam model with homogeneity and symmetry (at the sample mean)

Eq.[a]	US	MX	CD	Exp
US	-0.8468*	-0.1910*	-0.0262*	1.0640*
	(0.0678)	(0.0531)	(0.0107)	(0.1019)
MX	-0.2843	-0.5954*	0.0107	0.8690*
	(0.1553)	(0.1236)	(0.0267)	(0.2335)
CD	-0.3061	0.0460	-0.6938*	0.9538*
	(0.2418)	(0.2093)	(0.1836)	(0.3569)

[a] Abbreviations are the same with the previous table.
* indicates a coefficient that is significantly different from zero at the 5% significance level ($p = 0.05$). Estimates in parentheses are standard errors.

Table 8. Compensated price elasticities for restricted Rotterdam model with homogeneity and symmetry (at the sample mean)

Eq.[a]	US	MX	CD
US	-0.1530*	0.1311*	0.0219*
	(0.0354)	(0.0369)	(0.0097)
MX	0.2824*	-0.3323*	0.0500*
	(0.0796)	(0.0873)	(0.0251)
CD	0.3159*	0.3347*	-0.6507*
	(0.1405)	(0.1678)	(0.1816)

[a] Abbreviations are the same with the previous table.
* indicates a coefficient that is significantly different from zero at the 5% significance level ($p = 0.05$). Estimates in parentheses are standard errors.

Table 9. Parameter estimates of restricted FD/AIDS with homogeneity and symmetry

Eq.[a]	US	MX	CD	Exp
US	0.0852*	-0.0819*	-0.0033	0.3279*
	(0.0174)	(0.0182)	(0.0051)	(0.0451)
MX		0.0699*	0.0119	-0.3472*
		(0.0202)	(0.0063)	(0.0483)
CD			-0.0086	0.0193
			(0.0068)	(0.0118)

[a] Abbreviations are the same with the previous table.
* indicates a coefficient that is significantly different from zero at the 5% significance level ($p = 0.05$). Estimates in parentheses are standard errors.

Table 10. Uncompensated price and expenditure elasticity for restricted FD/AIDS with homogeneity and symmetry (at the sample mean)

Eq.[a]	US	MX	CD	Exp
US	-1.1967*	-0.2802*	-0.0278*	1.5047*
	(0.0511)	(0.0361)	(0.0031)	(0.0694)
MX	0.4704*	-0.4239*	0.0901*	-0.1365
	(0.1157)	(0.0840)	(0.0071)	(0.1580)
CD	-0.3533	0.1352	-1.2111*	1.4292*
	(0.1990)	(0.1630)	(0.1530)	(0.2626)

[a] Abbreviations are the same with the previous table.
* indicates a coefficient that is significantly different from zero at the 5% significance level ($p = 0.05$). Estimates in parentheses are standard errors.

Table 11. Compensated price elasticity for restricted FD/AIDS with homogeneity and symmetry (at the sample mean)

Eq.[a]	US	MX	CD
US	-0.2192*	0.1794*	0.0397*
	(0.0268)	(0.0281)	(0.0079)
MX	0.3816*	-0.4656*	0.0840*
	(0.0597)	(0.0662)	(0.0206)
CD	0.5752*	0.5717*	-1.1469*
	(0.1146)	(0.1403)	(0.1521)

[a] Abbreviations are the same with the previous table.
* indicates a coefficient that is significantly different from zero at the 5% significance level ($p = 0.05$). Estimates in parentheses are standard errors.

References

Alston, J.M., C.A. Carter, R. Green, and D. Pick. "Whither Armington Trade Models?" American Journal of Agricultural Economics 72 (1990): 455-467.

Alston, J.M. and J.A. Chalfant, and N.E. Piggott. "Estimating and Testing the Compensated Double-Log Demand Model." Applied Economics 34 (2002): 1177-1186.

Armington, P.S. "A Theory of Demand for Products Distinguished by Place of Production." IMF Staff Papers 16 (1969): 159-177.

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

Florida Tomato Committee. Tomato 101. Orlando, FL, 2002. http://www.floridatomatoes.org/education.htm.

Kmenta, J. *Elements of Econometrics*. 2nd ed. Ann Arbor: University of Michigan Press, 1997.

LaFrance, J.T. "When is Expenditure "Exogenous" in Separable Demand Models?" Western Journal of Agricultural Economics 16 (1991): 49-62.

Theil, H. The System-wide Approach to Microeconomics. Chicago: University of Chicago Press, 1980.

U.S. Department of Agriculture, Economic Research Service. ERS/USDA Briefing Room – Tomatoes: Background. Washington, DC, December 2000. http://www.ers.usda.gov/briefing/tomatoes/background.htm.

U.S. Department of Agriculture, Economic Research Service. U.S. Tomato Statistics. Washington, DC, March 2003. http://usda.mannlib.cornell.edu/data-sets/specialty/92010/.

U.S. Department of Agriculture, Economic Research Service. Vegetables and Melons Situation and Outlook Yearbook. VGS-2003. Washington, DC, July 2003. http://jan.mannlib.cornell.edu/reports/erssor/specialty/vgs-bb/2003/vgs2003.pdf.

U.S. Department of Agriculture, Foreign Agricultural Service. U.S. Trade Imports – FATUS Commodity Aggregations. Washington, DC, January 2004. http://www.fas.usda.gov/ustrade/USTImFatus.asp?QI=.

Weatherspoon, D.D. and J.L. Seale, Jr. "Do the Japanese Discriminate against Australian Beef Imports?: Evidence from the Differential Approach." Journal of Agricultural and Applied Economics 27 (1995): 536-543.

Winters, L.A. “Separability and the Specification of Foreign Trade Functions.” Journal of International Economics 17 (1984): 239-263.

Zellner, A. “An Efficient Method of Estimating Seemingly Unrelated Regressions and Tests for Aggregation Bias.” Journal of the American Statistical Association 57 (1962): 348-368.