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February 1989

AFPC Staff Report 89-3

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This paper was prepared for presentation at the Southern Agricultural Economic Association Annual Conference on February 3-8, 1989, Nashville, TN. The research was partially supported by a cooperative agreement, project No. 58-3AEK-8-00019, between Texas Agricultural Experiment Station, Texas A&M University and the CED/ERS/USDA.

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Abstract

The Armington procedure has become increasingly popular in agricultural trade analyses. However, some arguments have arisen concerning the relevance of using the procedure for agricultural trade analyses. This study examines the assumptions commonly made when using the Armington procedure. The results of this research on rice trade suggest that the Armington procedure needs some modification in modelling agricultural trade and that the assumptions of the single constant elasticity of substitution and homotheticity, in particular, may not be appropriate for analyzing agricultural trade. Further, results of a modified Armington procedure indicate that rice exports are highly competitive and that change in market shares of individual exporters are not independent of change in budget expenditure allocated to imports.

Key words: Armington procedure, trade, rice, multi-CES, homotheticity.

Introduction

Several researchers have employed the procedure developed by Armington for agricultural trade modeling. An early application of this procedure was that of Grennes *et al.* in a study of world grain trade. Thompson argued in his extensive research review that the Armington procedure (AP) is a very promising approach for agricultural trade analyses. Since then the AP has been widely used in studies of agricultural trade. Honma and Heady (subsequently expressed as H/H) developed a wheat trade model using the Armington procedure. Two years later Babula used the procedure in a study of world wheat, corn, and cotton markets, and Figueroa and Webb (subsequently F/W) also employed it for their wheat and corn trade analyses. The three studies cited are all based on time-series data. However, their model specifications and empirical results are quite different from one study to another (Table 1).

There are some important modeling issues involved in the use of the AP for agricultural trade analyses. The procedure does not account for an importer's domestic production, an important determinant of trade flows in agriculture. Another issue concerns that the single constant elasticity of substitution (CES) assumption, a major assumption in the AP, may not be accurate for agricultural trade analyses (Thompson). Further, shares of individual exporters may not be homothetic and the preferences of importers for products originating from different suppliers may be a critical factor for determing market-shares (Winters; Carter *et al.* ; and Ito *et al.*).

In this paper, the applicability of the AP for agricultural trade analyses is reviewed, and specification problems for modeling agricultural trade are discussed. Finally, the adequacy of the original assumption of the AP are tested for a specific example of agricultural trade employing an alternative approach that retains the basic concept of the AP.

Evaluation of the Armington procedure for Agricultural Trade Studies

Armington attempts to differentiate products from different suppliers in a market. He employs a two stage procedure, assuming that, at the first stage, a "buyer" decides on the total volume to purchase, and at the second stage, allocates portions of the total volume to individual suppliers in order to minimize the costs. For the first stage equation, he specifies the total demand for both foreign and domestic products as the dependent variable. Assuming that a "buyer" maximizes his utility, U, subject to available income, his utility would be:

$$(1)\quad \text{Max}\quad U = (Q_1,\ Q_2\ \ldots\ldots\ Q_n),$$
$$\text{subject to}\quad Y = \Sigma Q_i P_i,$$

where, Q_i is the i-th good or market consisting of a group of products, P_i is price index in the i-th market, and Y is income. Forming a Lagrangean equation and solving the first order conditions, a Marshallian demand function for Q_i can be derived:

$$(2)\quad Q_i = f(Y,\ P_1,\ P_2,\ \ldots.\ P_n).$$

For the second stage equation, Armington makes two major assumptions: 1) the elasticity of substitution is constant (CES) regardless of the share of a product; 2) there is a single elasticity of substitution between any pair of products in the group. The two assumptions, which are together regarded as the single CES assumption, allow us to reduce the number of coefficients to be estimated and make the estimation process easier. Under these assumptions, Armington specifies the second stage equations as follows:[1]

$$(3) \quad q_{ij} = b_{ij}^{\sigma_i} Q_i (P_{ij}/P_i)^{-\sigma_i}, \text{ or}$$

$$(4) \quad q_{ij}/Q_i = b_{ij}^{\sigma_i} (P_{ij}/P_i)^{-\sigma_i},$$

where, q_{ij} is a product from the j-th supplier to the i-th market, σ_i is the single constant elasticity of substitution for the products in the i-th market, P_{ij} is the price of q_{ij}, and q_{ij}/Q_i is the market share of the j-th product in the i-th market.

Armington originally developed the procedure to analyze trade in products such as chemicals under an assumption that there are no major trade restrictions. In his example, 20 suppliers of chemicals including the domestic suppliers sell in a market with no barriers to the imported products. In other words, the "buyer" or the importing country only considers relative prices among the products from different suppliers. This restriction on the importer's behavior with respect to imported products leads to some technical problems in applying the AP for agricultural trade analyses.

First, for Armington's first stage equation, several problems specific to agricultural trade need to be addressed. For chemical products, which Armington used as an example, trade restrictions are generally "technical" with some cases of tariffs, and import quotas on chemical products are very few (U.S. Trade Representative, 1987 and 1988). Armington assumes that "import demands are not residual demands depending upon domestic supply functions" (p.163). This is the reason why Armington theoretically constructed a demand function for all chemical products

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[1] See Armington (pp. 172-173) for detailed mathematical derivation process for Equation(3).

including both domestic and imported products in the first stage. (See Armington, pp. 161-164).

In agriculture, however, the trade situation is different. Agricultural trade is often controlled by governments in an effort to stabilize domestic prices, reduce dependency on foreign products, reduce foreign debts, or protect domestic producers.[2] In a world of this nature, trade ought to be considered a residual which contradicts Armington's original assumption. This assumption is reasonable only if free trade obtains in the importing country. This condition is not found in most agricultural markets, particularly if the government sets the quantity of imports allowed with import quotas. Therefore, an approach that does not recognize the residual nature of agricultural trade may not be appropriate. To incorporate the residual nature of agricultural trade, H/H and F/W specified a total import demand equation with domestic production as an explanatory variable in their first stage equations (Table 1).

The second problem of the Armington procedure for agricultural trade analyses is the specification of the second stage equation. For example, using the double log form and linearizing Equation(3), a quantity dependent equation, gives the following:

$$(5)\ \ln(q_{ij}) = \sigma_i \ln b_{ij} + \ln(Q_i) - \sigma_i \ln(P_{ij}/P_i).$$

In an econometrically estimated equation, the first two terms on the right-hand side, $\sigma_i \ln b_{ij} + \ln(Q_i)$, are estimated as the intercept, while the price coefficient, $(-)\sigma_i$, of the third term on the right-hand side is interpreted as the single CES. The total volume of imports can fluctuate

[2] There is an almost unlimited number of publications regarding protection of domestic agriculture and trade policies. For example, see Yeats, Bredahl *et al*, Johnson *et al*, McCalla and Josling, and Peterson.

due to variation of the importer's domestic production even if world prices remain at the same level. Accordingly, the fluctuation of import volume may not be explained by relative prices alone.[3]

Employing Equation(4), on the other hand, the problem above may be avoided. The dependent variable of the equation is expressed in the form of a market share. Using the double log form, the equation is expressed as follows:

$$(6)\ \ln(q_{ij}/Q_i) = \sigma_i \ln b_{ij} - \sigma_i \ln(P_{ij}/P_i).$$

This specification avoids the problem of variations in import quantities arising from a fluctuation of an importer's domestic production. Although total imports fluctuate, it is not unreasonable to assume that the share of each exporter would not change unless relative prices change among the imported products. Thus, a market-share dependent equation for the second stage should be more realistic than the quantity dependent equation. H/H[4] and Babula[5] used the quantity dependent equation and found mostly insignificant coefficients for their price variables, while F/W used the market-share dependent equation and found the estimated coefficients generally significant (Table 2). If the data for the analysis are solely

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[3] Hickman and Lau developed an Arminton specification using Taylor's series expansion:

$$q_{ij} = b_{ij}^{o} Q_i - \sigma_i q_{ij}^{o}(P_{ij} - P_i),$$

where the definitions of the variables are the same as those in Equation(5) except for that b_{ij}^{o} and q_{ij}^{o} are b_{ij} and q_{ij} in the base year, respectively. In this specification, however, the dependent variable is also expressed in quantity. Thus, it may be possible that the Hickman-Lau specification causes the same kind of problem as Equation(5).

[4] H/H employed the specification developed by Hickman and Lau.

[5] Babula in particular used the volume of total imports as an explanatory variable in the second stage based on the specification of Equation(3).

cross-sectional (with no time-series data), there should not be much difference between the results from using either quantity dependent or market-share dependent specification.

The third problem concerns the utilization of time-series data for the second stage Armington equation. Armington's original specification is implied to apply for cross-sectional data. Because the number of exporters of a specific agricultural commodity is generally small and because data availability for each exporter are also limited, there may not be enough observations in the cross-sectional data. Thus, a cross-sectional econometric estimate for agricultural trade is practically impossible. Hickman and Lau employed Taylor series expansion and used the AP for cross-sectional time-series analyses. In their study, they pooled data for each importer and estimated Armington's single CES for each one of importing regions using trade flow data over time. F/W pooled cross-sectional and time-series data and employed dummy variables to adjust intercepts for individual exporters.[6] By pooling data, it is technically possible to increase the number of observations. However, due to competition among the individual exporters who aim to expand their shares in an importing region, it is likely that behaviors of the exporters are not mutually independent from one another. Accordingly, the error terms among equations estimated separately using OLS for individual suppliers may be correlated. In such cases, data cannot be pooled and used to estimate a single equation. In addition, it is appropriate to estimate the single coefficient out of the pooled data *only if* the estimated coefficients in the separate equations are the "same" (Kmenta, p. 518). Estimated coefficients using pooled data may then be inefficient. To solve

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[6] More recently, Haniotis and Ames used the same technique for a study on soybean exports to the EC.

this problem, it is necessary to employ the seemingly unrelated regressions (SUR) originally developed by Zellner.[7]

Fourth, some writers have criticized the single CES assumption (Branson; and Thompson). Thompson notes that "there seems to be a logical inconsistency between assuming a commodity is differentiated by country of origin and then assuming the same parameters," (p. 44). None of the previous studies cited above tested whether or not imposing the assumption of a single CES is appropriate for agricultural trade analyses. If results from the test showed that the estimated price coefficients are different from one another, it can be concluded that the assumption of the single CES of the AP may not be appropriate.

Finally, the AP retains an assumption of homotheticity, which implies that export shares of individual suppliers are independent of overall level of budget allocated to the imports in a specific importing region. Winters tested homotheticity using trade data considering only manufactured goods. Carter *et al.*, using cotton and wheat trade data, also tested the homotheticity assumption. Both Winters and Carter *et al.* applied AIDS model and rejected a hypothesis of homotheticity, stating that Armington's assumption is not acceptable. Ito *et al.* found that different importers have distinct preferences for rice from different sources so that shares varied as budget allocated to imported rice changed. Accordingly, it is important to test the homotheticity condition in analysis of agricultural trade.

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[7] H/H specified a single time-series equation for each of five suppliers to an importer and used SUR. In their model, they restricted the coefficient of the price variable in each equation to be identical. This was done because of Armington's single CES assumption.

A Modified Approach for Applying the Armington Procedure to Agricultural Trade

To solve the problems described above, an alternative method of applying the Armington procedure for agricultural trade is proposed. First, the dependent variable in Armington's first stage equation is the total demand including products from both domestic and foreign suppliers. Before moving to the second stage, it is necessary to incorporate the residual nature of agricultural imports in agricultural trade. Thus, another equation is inserted in order to explain the total *import* demand, and it is specified as the following identity:

$$(7) \quad Q_i^* = D_i + H_i - S_i - H_{i,-1},$$

where, Q_i^* is total import demand, D_i is domestic demand, H_i is ending stocks, and S_i is domestic production. Variables D, H, and S have to be estimated individually in stochastic structural equations, because they are considered to be endogenous.[8] This specification avoids the problem of accounting for the influence of domestic production on imports in a stochastic equation.

For Armington's second stage equation (or trade-flow equation), in particular, it might be better to use market-share as the dependent variable rather than quantity, although this proposition is testable. A market-share equation for the second stage is particularly useful in analyzing the competitiveness of one supplier relative to others.

Assuming that a single-CES does not hold, it is necessary to derive an equation for estimating multi-CES. An alternative specification for

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[8] This specification is a part of the nonspatial price equilibrium procedure according to the Thompson's terminology (Thompson). In developing a world model, this procedure has been widely employed (Devadoss *et al.* ; and Meyers *et al.*).

estimating multi-CES can be derived based on Equation(4) by replacing σ_i with σ_{ij}:

$$(8)\quad (q_{ij}/Q_i^*)_t = b_{ij}^{\sigma_{ij}}(P_{ij}/P_i)_t^{-\sigma_{ij}},$$
$$j = 1, 2, \ldots m; \quad t = 1, 2, \ldots T.$$

The partial derivative of q_{ij}/Q_i^* with respect to P_{ij}/P_i is:

$$(9)\quad \partial(q_{ij}/Q_i^*)/\partial(P_{ij}/P_i) = b_{ij}^{\sigma_{ij}}(-\sigma_{ij})(P_{ij}/P_i)_t^{-\sigma_{ij}-1}.$$

Therefore, a constant elasticity of substitution (CES) is:

$$(10)\quad \{\partial(q_{ij}/Q_i^*)/\partial(P_{ij}/P_i)\}\ \{(P_{ij}/P_i)/(q_{ij}/Q_i^*)\} = -\sigma_{ij}.$$

The $(-)\sigma_{ij}$ is a CES of the products from the j-th exporter to the i-th market relative to the total of the good's market. Thus, the number of σ_{ij}'s is the number of exporters of the commodity to the region. Using a SUR without restriction on price coefficients, it is possible to estimate a CES for each product originating from a specific exporting country.

Finally, to incorporate a factor for testing homotheticity, *i.e.,* independence of change in share of an exporter from total level of budget allocated to imports, it is necessary to modify Equation(7). To do this, the b_{ij} in Equation(7) is replaced by B_{ij}:

$$(11)\quad (q_{ij}/Q_i^*)_t = B_{ij}^{\sigma_{ij}}(P_{ij}/P_i)_t^{-\sigma_{ij}},$$

where,

$$(12)\quad B_{ij} = b_{ij}EXP(\theta_{ij}\ln(X_i/P^*)).$$

The θ_{ij} in Equation(12) is a coefficient of expenditure for the j-th supplier to the i-th importing region, X_i is the total budget expenditure allocated to products from all suppliers to the i-th importing region, and P^* is a price index to serve as a filter in order to avoid money illusion for X_i.

Linearizing Equation(12) using the log form, it is expressed as follows:

(13) $\ln(q_{ij}/Q_i^*)_t = \sigma_{ij}\ln b_{ij} - \sigma_{ij}\ln(P_{ij}/P_i)_t + \sigma_{ij}\theta_{ij}\ln(X_i/P^*)_t,$

or,

(14) $\ln(q_{ij}/Q_i^*)_t = \sigma_{ij}\ln b_{ij} - \sigma_{ij}\ln(P_{ij}/P_i)_t + \beta_{ij}\ln(X_i/P^*)_t,$

where β_{ij} (= $\sigma_{ij}\theta_{ij}$) is an elasticity of budget expenditure for products from the j-th country to the i-th importing region. Thus, there exist the same number of estimated β as the number of exporters. If all β_{ij}'s are found to be not significantly different from zero, it is concluded that rice imports are homothetic and that imports of rice from a specific supplier are independent of the level of budget allocated to rice imported. If, on the other hand, at least one of β_{ij}'s is found to be different from zero, homotheticity does not hold. A positive estimated β_{ij} indicates that market share of the j-th supplier increases as the allocated budget in the i-th importing region increases. This can also be interpreted to mean that the importing region tends to consume more the products from the j-th country at the expense of other suppliers' shares as the allocated budget to imports increases. The larger the absolute value of β, the more elastic the preference for the products in the importing region.

Data

Trade flow data on rice were collected from the *Commodity Trade Statistics* (United Nations) for twenty-five years, 1962-1986, based on calendar years. Imports by an individual country not only vary but are often zero in certain years. This makes econometric analysis more difficult. As a result, all the importing countries and regions were aggregated into one region. Exporters are categorized into seven regions: Thailand, the U.S., Argentina, Australia, Burma, Italy, and Pakistan.[9] Exporters that have not reported their data to the United Nations were excluded.[10] Shipments by the U.S. Government under concessional government-financed programs such as P.L. 480, foreign donations (Section 416), and AID mutual security programs are excluded, because actual prices for these types of shipments deviate considerably from the market prices. These U.S. Government shipment data were collected from *FATUS* (U.S. Department of Agriculture). Export prices were calculated from the total export value divided by total quantity for each exporting country and all expressed in U.S. dollars. Data of budget allocated to imports are not available; therefore, X_i was approximated by total expenditure for imports of the products from individual suppliers. Consumer price index in the U.S. was employed as price index, P^*, because X_i was expressed in U.S. dollars.

Empirical Results of Estimating CES

In this section, the results of testing the specification of original and modified Armington's second stage equations are reported. First, the relevance of using pooled data under the single CES assumption was tested by running a regression for equations with market-share as the dependent variable. According to a test of correlation coefficients among the error terms generated from ordinary least square (OLS) equations for individual

[9] To avoid the singularity problem in the SUR, an equation for Argentina, which is the smallest rice exporter among the seven nations, is deleted.

[10] The People's Republic of China, a major rice exporter, is excluded because of this problem.

exporters, serious correlations were observed. In addition, the estimated price coefficients are not the same but statistically different from one another in the equations for individual exporters.[11] These results suggest that the coefficients estimated by procedures such as pooling data or using OLS are inefficient.

Second, quantity dependent and market-share dependent equations under the original Armington assumption were compared. The quantity dependent equation is the double-log form from Equation(5) -- (Model I), and the market-share equation is the double-log model from Equation(6) -- (Model II). The results are reported in Table 3. There is no formal statistical test to decide superiority between the two specifications because the dependent variables are not identical. However, judging from the estimated R-square in each equation in both models, it appears that the market-share dependent specification in Model II is superior to the others. The R-squares in Model II are mostly greater than those in Model I, while two of the R-squares in Model I are negative.[12]

Third, the appropriateness of the single-CES and homotheticity assumptions is tested. The results above indicate that a market-share dependent specification appeares to be superior to the quantity dependent specification. Therefore, the system of equations in Model II, the original Armington single-CES and homotheticity assumptions with market-share dependent specification, was compared with that of Model III, multi-CES and non-homotheticity assumptions based on Equation(14). This is to jointly test whether or not a specification based on the multi-CES and

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[11] Results of the estimated OLS equations can be obtained from the authors.

[12] The negative R-squares in Model I may be due to the restriction of price coefficients being identical for all suppliers.

non-homotheticity assumptions is statistically superior to a specification based on the single-CES and homotheticity assumptions. A test of a set of linear restrictions was performed using F-statistics estimated from the Lagrange multiplier (LM) (Judge *et al.*, 1982, pp. 326-328 and Judge *et al.*, 1985, pp. 472-477). The set of restrictions are expressed by:

(15) $\mathbf{RB} - \mathbf{r} = \mathbf{0}.$

Thus, the hypothesis is:

Ho: $\mathbf{RB} - \mathbf{r} = \mathbf{0},$

Ha: $\mathbf{RB} - \mathbf{r} \neq \mathbf{0}.$

If Ho is rejected, it is concluded that restrictions under the assumption of single-CES and homotheticity are inappropriate. The system of equations in Model II is such that restrictions on price coefficients (single-CES) and budget coefficients (0 under homotheticity) are imposed. On the other hand, the Model III system has no restrictions.

The Lagrange multiplier test is expressed as follows:

(16) $\lambda = A_J \,/\, B_{MT-K} \sim F_{J,MT-K}$,

where,

$$A_J = \{(\mathbf{y} - \mathbf{X\hat{B}^*})'(\Sigma^{-1} \times \mathbf{I})(\mathbf{y} - \mathbf{X\hat{B}^*}) - (\mathbf{y} - \mathbf{X\hat{B}})'(\Sigma^{-1} \times \mathbf{I})(\mathbf{y} - \mathbf{X\hat{B}})\} \,/\, J,$$

$$B_{MT-K} = (\mathbf{y} - \mathbf{X\hat{B}})'(\Sigma^{-1} \times \mathbf{I})(\mathbf{y} - \mathbf{X\hat{B}}) \,/\, (MT - K),$$

$\mathbf{\hat{B}^*}$ represents estimated coefficients under restrictions, and $\mathbf{\hat{B}}$ represents the estimated coefficients under no restrictions. Σ is the covariance matrix, J is a number of restrictions, M is a number of equations in each system, K is a number of explanatory variables including the intercepts in the system with no restrictions, and T is the number of observations in each equation.

The Lagrange multiplier test resulted in an estimated F-value being equal to 3.451, which is greater than 2.18, the critical value at the 1% significance level for degrees of freedom of v_1=12 and v_2=121 to ∞ (Table 4). Accordingly, Armington's original assumptions of the single-CES and homotheticity are statistically rejected at 1% significance level. This indicates that the system of equations in Model III is statistically superior to those in Model II. In fact, the estimated price coefficients in Model III vary among individual exporters and are all statistically significant. Further, coefficients of budget expenditures were significant for two exporters; a negative coefficient for Burma and positive one for Pakistan. This indicates that market shares of individual exporters are not always independent of change in level of budget expenditures and that an assumption of homotheticity in the AP may be erroneous. These results suggest that Armington's original assumptions of the single-CES and homotheticity may not be appropriate for this particular market. On the other hand, it is possible to modify the original Armington specification as illustrated in this case by Model III.

Conclusion

The Armington procedure is becoming more popular for agricultural trade analyses. In this paper, Armington's original procedure and three recent empirical studies using the procedure for agricultural trade are evaluated. Then, the relevance of Armington's assumptions is examined, and a modified approach is proposed for agricultural trade analyses.

The results of this research suggest that direct application of the procedure for agricultural trade analyses may not be appropriate. The empirical results show that the assumption of the single CES is not consistent with the data for world rice markets. In addition, homotheticity is not an appropriate assumption for this market. The results in this research are basically consistent with those found by Winters and Carter *et al*. However, these results do not necessarily imply that the Armington's basic concept should be totally rejected. Rather, the Armington procedure can be a powerful method to analyze agricultural trade, if it is properly modified.

The alternative method proposed includes the following modifications: 1) Replace Armington's first demand equation with a total import demand equation estimated by an identity derived from a structural model in order to account for the influence of domestic production on imports; 2) Estimate the second stage equation using market-share as the dependent variable instead of quantity to evaluate the effects of changes in relative prices and expenditures for quantities imported, and 3) Test the assumptions of the single-CES and homotheticity and adjust the specification of the model if the assumptions are rejected.

Implications for Rice Trade

The estimated coefficients in this study for rice trade indicate that an elasticity of substitution between products from a specific exporter and the other exporters vary among themselves. The estimated constant elasticity of substitution for Australia is the largest, while those for Burma and Italy are at the smallest. The relatively small constant elasticity of substitution for Italy at -1.284 may be due to the influence of the Common Agricultural Policies (CAP) in the European Community (EC). Under the CAP, trade within the community is promoted by using "a common external tariff applied to trade with countries outside the region while eliminating all tariffs within the community" (Peterson). Rice exports from Italy to other EC member countries accounted for approximately 40% of total rice exports from that country in 1986. The constant elasticity of substitution estimated for U.S. rice exports is the third smallest at -1.519. This implies that the U.S. faces less secure markets than in the case of Italy.

Overall, however, the estimated constant elasticities of substitution for each rice exporter are generally greater than unity. This indicates that import demand for rice products from a specific country is very sensitive to relative prices. Accordingly, it is clear that rice export markets are highly competitive. In this context, rice exporters must keep prices in line with prevailing world prices in order to maintain competitiveness in world rice markets. Further, a policy such as the marketing loan, introduced with the 1985 Farm Bill aimed at decreasing export prices and insulating U.S. prices from changes in exchange rates or other domestic economic variables, is an effective tool to recover the U.S. share in the world markets.

Finally, the estimated coefficients for the budget expenditure varied among suppliers. Two exporters, Thailand and Burma, had negative coefficients, while the others had positive coefficients. This may reflect the situation that rice from Thailand and Burma is generally considered to be inferior to rice from other exporters. On the other hand, the coefficient for Pakistan is the largest, and it is positive. Rice from Pakistan is mainly aromatic rice called "basmati," and is more expensive than rice from the other exporters. These results strongly suggest that rice importers are selective among the products from different suppliers as they change their budgets allocated to rice imported. Therefore, it is important for exporters to improve and maintain the quality of their products.

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Table 1. Description of first and second stage equations in previous empirical researches applying the Armington procedure.

	Honma/Heady (1984)	Babula (1986)	Figueroa/Webb (1986)
1st Stage Estimation			
Dependent variable	Total imports (per capita)	Total imports	Total imports
Inclusion of domestic prodn. as explanatory var.	Yes	No	Yes
Other independent variables	Wheat price Corn price Income Ending stocks Gov. imports Dummy var.	Wheat price Corn price GDP Oil price Lag dept. var.	Wheat price Corn price GNP CPI Dummy var.
2nd Stage Estimation			
Time-series/ Cross-sectional	T-S with SUR	T-S	C-S with T-S
Dependent variable	q_{ij}	q_{ij}	q_{ij}/Q_i
Price variable	$P_{ij} - P_i$	P_{ij}/P_i	P_{ij}/P_i
Other independent variables	Time trend Dummy var.	Total imports Time trend Ship service supply Oil price index	Intercept dummies
# of importers	10	6	8
# of exporters	5	U.S. & ROW	6

Table 2. The previously estimated single CES for wheat in individual importing regions.

Honma/Heady (1984)		Babula (1986)		Figueroa/Webb (1986)	
EC-6	--	EC-10	-3.43 (0.69)	EC	-0.46 (1.33)
EC-3	-0.7538 (2.48)	Brazil	-4.13 (0.70)	Egypt	4.36 (4.37)
Japan	--	Japan	-1.13 (0.83)	Japan	-1.75 (2.26)
RODC[1]	-0.4768 (6.01)	S.America[2]	-0.16 (0.10)	S.Korea	-7.68 (5.92)
NIC[3]	--	N.Africa	-8.08 (1.79)	Taiwan	-8.74 (2.37)
OPEC	-0.6075 (2.78)	ROW	-2.17 (1.25)	USSR	-0.374 (0.30)
ROLDS[4]	-2.1214 (4.17)			Mexico	-6.63 (1.93)
China	--			China	-4.06 (3.78)
USSR	--				
E. Europe	-0.3163 (3.15)				

( ) = t-values.
-- indicates no estimated price coefficient reported.
1:Switzerland, Portugal, and Israel.
2:Excluding Brazil.
3:Brazil, Mexico, and S. Korea.
4:India, Pakistan, Egypt, Morocco, Peru, and the Philippines.

Table 3. Comparison of different specifications among original and modified Armington procedures.[1]

Method:	Original Armington procedure				Modified Armington procedure		
Assumption:	Single-CES and homotheticity				Multi-CES and non-homotheticity		
Model:	Model I		Model II		Model III		
Dep. var.:	Quantity		Market-share		Market-share		
	R^2	$-\sigma_i$	R^2	$-\sigma_i$	R^2	$-\sigma_{ij}$	β_{ij}
Country:							
Thailand	0.282	-1.683 (0.126)	0.210	-1.689 (0.111)	0.264	-1.586 (0.320)	-0.100 (0.094)
U.S.	0.396	" (")	0.553	" (")	0.555	-1.519 (0.304)	0.082 (0.145)
Australia	0.315	" (")	0.546	" (")	0.576	-1.851 (0.340)	0.122 (0.185)
Burma	-0.067	" (")	0.182	" (")	0.444	-0.984 (0.489)	-1.291 (0.386)
Italy	-0.342	" (")	0.058	" (")	0.197	-1.284 (0.288)	0.388 (0.261)
Pakistan	0.473	" (")	0.624	" (")	0.724	-1.726 (0.301)	0.546 (0.229)

Standard errors are in parentheses.
Those " indicate being identical to the number above.
1: The seemingly unrelated regression (SUR) is used in each model.

Table 4. Results of hypothesis testing for Armington's original assumptions.

$\Sigma_j SSE_{ij}$		J	MT-K	F-value	$F_{1\%,12,\infty}$
Model II	Model III				
31.966	24.333	12	132	3.451	2.18