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Modeling international trade flows and market shares for agricultural commodities: a modified Armington procedure for rice

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ABSTRACT

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The Armington procedure (AP) has become increasingly popular in agricultural trade analyses. However, some arguments have arisen concerning the relevance of using the procedure for such analyses. This study examines the assumptions commonly made when using the Armington procedure and suggests modifications for agricultural trade analyses. Results from models utilizing rice-trade data suggest that the assumptions of the single constant elasticity, in particular, may not be appropriate for analyzing agricultural trade. These results also suggest that, with proper modifications, the AP can be applied to agricultural trade. Further, results of a modified Armington procedure indicate that trade in rice exports is highly competitive and that changes in market shares of individual exporters are not independent of changes in budget expenditure allocated to imports.

Introduction

Several researchers have employed the procedure developed by Armington (1969) for agricultural-trade modeling. The procedure is designed to determine trade flows explicitly. An early application of this procedure was that of Grennes et al. (1973) in a study of world grain trade. Thompson (1981) argued in his extensive research review that the Armington procedure (AP) is a very promising approach for agricultural trade analyses, because it allows estimation of parameters reflecting the behavior of importers faced

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with an array of similar products differentiated by origin. Since then, the AP has been widely used in studies of agricultural trade. Honma and Heady (1984) (subsequently expressed as HH) developed a wheat-trade model using the Armington procedure. Two years later, Babula (1986) used the procedure in a study of world wheat, corn, and cotton markets; Figueroa and Webb (1986) (subsequently FW) also employed it for their wheat- and corn-

TABLE 1

Description of first- and second-stage equations in three previous empirical studies 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 production as explanatory variable	Yes	No	Yes
Other independent variables	Wheat price Corn price Income Ending stocks Government imports Dummy variables	Wheat price Corn price GDP Oil price Lag dependent variables	Wheat price Corn price GNP CPI Dummy variables
2nd-stage estimation			
Time-series (T-S)/ Cross-sectional (C-S)	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 variables	Total imports Time trend Ship service supply Oil price index	Intercept dummies
# of importers	10	6	8
# of exporters	5	U.S.A. & ROW	6

SUR implies seemingly unrelated regression.
ROW implies rest of the world.

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 limitations involved in the use of the AP for agricultural trade analyses. Domestic production is often an important determinant of trade flows in agriculture because governments frequently introduce policy measures that have the effect of protecting the market share of domestic producers. In the original AP, this aspect of agricultural trade is not specifically included, although it is conceptually possible to include domestic production as a source of supply in the second stage. Another limitation is the assumption of a single constant elasticity of substitution (CES), a major assumption in the AP, which may not be accurate for agricultural-trade analyses (Thompson, 1981). 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 determining market shares (Winters, 1984; Ito et al., 1988; Alston et al., 1989).

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 assumptions of the second stage in 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-step 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 utility, U, subject to available income, the problem is to:

$$\text{Max } U = U(Q_1, Q_2, \ldots, Q_n)$$

subject to

$$Y = \Sigma_i Q_i P_i \tag{1}$$

where Q_i is the ith good or market consisting of a group of products, P_i is a price index for the ith 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:

$$Q_i = f(Y, P_1, P_2, \ldots, P_n) \tag{2}$$

For the second-stage equation, Armington makes two major assumptions: (1) the elasticity of substitution is constant 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 generalized CES form for Q_i:

$$Q_i = (\Sigma_j\, b_{ij} q_{ij}^{\;-\varrho_i})^{-1/\varrho_i} \quad (j = 1, 2, \ldots, m) \tag{3}$$

where $\Sigma_j\, b_{ij} = 1$, q_{ij} is a product from the jth supplier to the ith market, and ϱ_i is a constant for the ith market. Rewriting ϱ_i as $(1/\sigma_i - 1)$, he derives a CES demand function for q_{ij}:[1]

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

where σ_i is the constant elasticity of substitution for the products in the ith market, and P_{ij} is the price of q_{ij}. Equation (4) is expressed as a quantity-dependent equation. To specify the equation as a market-share-dependent equation, both sides are divided by Q_i:

$$q_{ij}/Q_i = b_{ij}^{\;\sigma_i}(P_{ij}/P_i)^{-\sigma_i} \tag{5}$$

Armington (1969) 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, twenty suppliers of chemicals including the domestic suppliers sell in a market with no major barriers to 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 telephone conversation with personnel in the Chemical Division, June 1988). Armington (1969) assumes that 'import demands are not residual demands depending upon domestic supply functions' (p. 163). This is the reason why Armington constructed a demand

[1] See Armington (1969, pp. 172 – 173) for the detailed mathematical derivation of equation (4).

function for all chemical products including both domestic and imported products in the first stage (see Armington, 1969, 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 mentioned in the preceding paragraph. 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, HH and FW specified a total import demand equation with domestic production as an explanatory variable in their first-stage equations (see 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 (4), a quantity-dependent equation, gives the following:

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

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 therm on the right-hand side is interpreted as the single CES. However, the total volume of imports can fluctuate 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 (5), 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

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[2] There is an almost unlimited number of publications regarding protection of domestic agriculture and trade policies. For example, see Yeats (1979), Bredahl et al. (1979), Johnson et al. (1985), McCalla and Josling (1985) and Peterson (1985).

[3] Hickman and Lau (1973) developed an Armington specification using Taylor's series expansion:

$$q_{ij} = b_{ij}^0 Q_i - \sigma_i q_{ij}^0 (P_{ij} - P_i),$$

where the definitions of the variables are the same as those in equation (6) except that b_{ij}^0 and q_{ij}^0 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 (6).

follows:

$$\ln(q_{ij}/Q_i) = \sigma_i \ln b_{ij} - \sigma_i \ln(P_{ij}/P_i) \tag{7}$$

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 preferable to the quantity-dependent equation. HH[4] and Babula[5] used the quantity-dependent equation and found mostly insignificant coefficients for their price variables, while FW used the market-share-dependent equation and found the estimated coefficients generally significant (Table 2). If the data for the analysis are solely cross-sectional (with no time-series data), there should not be much difference between the results from using either a quantity-dependent or a market-share-dependent specification. It is interesting to note the wide range of estimated elasticities of substitution for individual regions reported in Table 2. The great variation in these estimates suggests that the results in analyses of agricultural markets may be highly sensitive to the specification chosen.

The third problem concerns the utilization of time-series data for the second-stage Armington equation. Armington's original specification is derived for applications to cross-sectional data. Because the number of exporters of a specific agricultural commodity is generally small, and because data availability for each exporter is also limited, there may not be enough observations in the cross-sectional data. Furthermore, the use of time-series data allows inferences to be drawn concerning the reasons for changes in trade flows and market shares which occur over time. Cross-sectional data are useful in analyzing differences between countries but cannot capture temporal changes. Hickman and Lau (1973) employed Taylor's 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 importing region using trade-flow data over time. FW 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 impor-

[4] HH 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 (4).

[6] More recently, Haniotis and Ames (1988) used the same technique for a study on soybean exports to the EC.

TABLE 2

Estimated single CES for wheat in individual importing regions determined by three previous studies

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

( ) = *t*-values.
– indicates no estimated price coefficient reported.
[a] RODC: Switzerland, Portugal, and Israel.
[b] Excluding Brazil.
[c] NIC: Brazil, Mexico, and S. Korea.
[d] ROW, the rest of the world.
[e] ROLDS: India, Pakistan, Egypt, Morocco, Peru, and the Philippines.

ting region, it is likely that behaviors of the exporters are not mutually independent. Accordingly, the error terms among equations estimated separately using ordinary least squares (OLS) analysis 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, 1971, p. 518). Estimated coefficients using pooled data may then be inefficient. To solve this problem, it is necessary to employ the seemingly unrelated regressions (SUR) originally developed by Zellner (1963).[7]

Fourth, some writers have criticized the single CES assumption (Branson, 1972; and Thompson, 1981). Thompson (1981, p. 44) 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.' None of the previous trade studies cited above (HH, Babula, or FW) tested whether or not imposing the assumption of a single CES was appropriate for the markets they analyzed.

Finally, the AP retains an assumption of homotheticity, which implies that export shares of individual suppliers are independent of the overall level of budget allocated to the imports in a specific importing region. Winters (1984) tested homotheticity using trade data for manufactured goods only. Alston et al. (1989), using cotton- and wheat-trade data, also tested the homotheticity assumption. Both Winters and Alston et al. applied the almost ideal demand system (AIDS) model and rejected a hypothesis of homotheticity. Ito et al. (1988) 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 an analysis of particular agricultural markets.

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 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:

$$Q_i^* = D_i + H_i - S_i - H_{i,-1} \tag{8}$$

where Q_i^* total import demand, D_i domestic demand, H_i ending stocks, and S_i domestic production. Variables *D, H,* and *S* have to be estimated individually in stochastic structural equations, because they must be con-

[7] HH 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.

sidered endogenous.[8] This specification avoids the problem of accounting for the influence of domestic production on imports in a stochastic equation.

Regarding Armington's second-stage equation (or trade-flow equation), in particular, it can be concluded from the results of previous studies that market-share-dependent specifications tend to generate more statistically significant results than quantity-dependent specifications. This is examined empirically below for the case of rice.

Equation (7) includes the single CES assumption (only one CES for the entire market shared by all m suppliers) and the homotheticity assumption (no independent variable representing budget expenditure to explain the market share). In order to develop a multi-CES and nonhomothetic demand function for q_{ij}, consider equation (3). Suppose that importers operating in particular markets perceive differences in the characteristics of various products from individual exporting nations, and that these perceived differences cause them to treat products from individual exporters differently. Let these country-specific product characteristics be represented by the parameter γ_{ij}, reflecting the unique behavioral response of importer i with respect to products from exporter j. The parameter γ_{ij} may also include the differences in transportation costs between individual exporters and the importer. For this analysis, the parameter γ_{ij} is represented as an exponent for the variable q_{ij}. Thus, define $h_{ij} = q_{ij}^{\gamma_{ij}}$, where $\gamma_{ij} \neq 1$ for $\forall j$, and replace q_{ij} in equation (3) with h_{ij}:[9]

$$Q_i = (\Sigma_j \, b_{ij} h_{ij}^{-\varrho_i})^{-1/\varrho_i} \tag{9}$$

This allows the same derivation procedure to be followed as in the original approach, and euqation (4) can be rewritten as:

$$h_{ij} = b_{ij}^{\sigma_i} Q_i \, (P_{ij}/P_i)^{-\sigma_i} \tag{10}$$

Replacing h_{ij} by $q_{ij}^{\gamma_{ij}}$, Equation (10) would be:

$$q_{ij}^{\gamma_{ij}} = b_{ij}^{\sigma_i} Q_i \, (P_{ij}/P_i)^{-\sigma_i} \tag{11}$$

To specify the demand function as a market-share-dependent equation, move the exponent γ_{ij} to the right-hand side and divide both sides by Q_i:

$$q_{ij}/Q_i = b_{ij}^{\sigma_i/\gamma_{ij}} Q_i^{(1/\gamma_{ij}-1)} (P_{ij}/P_i)^{-\sigma_i/\gamma_{ij}} \tag{12}$$

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[8] This specification is a part of the nonspatial price equilibrium procedure according to Thompson's terminology (Thompson, 1981). In developing a world model, this procedure has been widely employed (Devadoss et al., 1986; Meyers et al., 1986).

[9] In the original Armington procedure, it is essentially assumed that $\gamma_{ij} = 1$ for $\forall j$.

In this equation, the market share is a function of total demand, Q_i, and the price ratio variable, P_{ij}/P_i. In the real world, the share of a specific product is not necessarily a function of total demand. If the quality of the product is inferior, that product's market share may decrease despite its low price, as buyers allocate more to the budget for a better class of products. It is also possible that the share of the low-quality product may increase because of its low price, if a buyer needs to purchase the same amount of the good with a reduced budget. Therefore, it is important to introduce a variable for the budget allocated to the good consisting of individual products in the market. Once a budget, V_i, for the products in the ith market is determined, actual expenditure can be less than or equal to this budget allocation. $V_i^{\mu_i}$ is the amount actually spent on good i (P_i times Q_i), where μ_i is a factor that allows less than the total budget allocation to be spent. If the budget is fully spent, μ_i is equal to 1; otherwise, it is smaller than 1, i.e. $0 < \mu_i \leq 1$. Accordingly, the relationship between V_i and Q_i is expressed as follows:

$$Q_i = V_i^{\mu_i}/P_i \qquad 0 < \mu_i \leq 1 \tag{13}$$

Replacing Q_i on the right-hand side in equation (12) with $V_i^{\mu_i}/P_i$ in equation (13) results in:

$$\begin{aligned} q_{ij}/Q_i &= b_{ij}^{\sigma_i/\gamma_{ij}}(V_i^{\mu_i}/P_i)^{(1/\gamma_{ij}-1)}\,(P_{ij}/P_i)^{-\sigma_i/\gamma_{ij}} \\ &= b_{ij}^{\sigma_i/\gamma_{ij}}\,V_i^{\mu_i(1/\gamma_{ij}-1)}\,P_{ij}^{-\sigma_i/\gamma_{ij}}\,(1/P_i)^{(-\sigma_i/\gamma_{ij}+1/\gamma_{ij}-1)} \end{aligned} \tag{14}$$

To simplify, equation (14) is rewritten:

$$q_{ij}/Q_i = b_{ij}^{\alpha_{ij}}V_i^{\theta_{ij}}P_{ij}^{\psi_{ij}}\,(1/P_i)^{\omega_{ij}} \tag{15}$$

where

$$\begin{aligned} \alpha_{ij} &= \sigma_i/\gamma_{ij} \\ \theta_{ij} &= \mu_i(1/\gamma_{ij} - 1) \\ \psi_{ij} &= -\sigma_i/\gamma_{ij} \\ \omega_{ij} &= -\sigma_i/\gamma_{ij} + 1/\gamma_{ij} - 1 \end{aligned}$$

Using the double-log form and adding a subscript for the time-series analysis, equation (15) can be rewritten:

$$\ln(q_{ij}/Q_i)_{\mathrm{t}} = \alpha_{ij}\ln b_{ij} + \theta_{ij}\ln(V_i)_{\mathrm{t}} - \psi_{ij}\ln(P_{ij})_{\mathrm{t}} + \omega_{ij}\ln(1P_i)_{\mathrm{t}} \tag{16}$$

Econometrically, it is possible to run a regression for equation (16). P_{ij} and P_i represent prices of the same kind of commodities, although P_{ij} is a price of differentiated product and P_i the weighted average price of products in the ith group. It is likely, therefore, that P_{ij} and $1/P_i$ will be highly correlated. To solve this problem, the two variables need to be combined and equation (16) respecified as:

$$\ln(q_{ij}/Q_i)_t = \alpha_{ij} \ln b_{ij} + \theta_{ij} \ln(V_i)_t + \sigma_{ij} \ln(P_{ij}/P_i)_t \tag{17}$$

Further, V_i is filtered by a price index, P^*, to avoid money illusion; Q_i is replaced by Q_i^* incorporating equation (8); and equation (17) is rewritten:

$$\ln(q_{ij}/Q_i^*)_t = \alpha_{ij} \ln b_{ij} + \beta_{ij} \ln(V_i/P^*)_t + \sigma_{ij} \ln(P_{ij}/P_i)_t \tag{18}$$

Equation (18) has to be specified for all suppliers to the *i*th market; $\{j = 1, 2, \ldots, m\}$. Because equation (18) is expressed in the double-log form, the coefficients, β_{ij} and σ_{ij}, are elasticities. The σ_{ij}, in particular, is regarded as a CES for the *j*th supplier to the *i*th market and expected to have a negative sign. Accordingly, the number of estimated CES is m, the number of suppliers.

The β_{ij}, on the other hand, is regarded as an elasticity of budget expenditure for products from the *j*th country to the *i*th market. Thus, there exists the same number of estimated β as the number of suppliers. 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 imported rice. If, on the other hand, at least one of β_{ij}'s is found to be different from zero, this implies that homotheticity may 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 market (or importing region) increases. This can also be interpreted to mean that the importing region tends to consume more of 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.

The analysis is for all the suppliers to a particular market in which the suppliers are competing with one another and behavior for a specific supplier is *not* independent of the behaviors for other suppliers in the market. Accordingly, the demand equations for all individual products from different suppliers have to be estimated simultaneously in a system. The seemingly unrelated regression analysis originally developed by Zellner is appropriate.

Finally, it is necessary to test whether or not the multi-CES and nonhomothetic approach is superior to the original AP. The system of equations in the original Armington single-CES and homotheticity assumptions with market-share-dependent specification expressed in equation (7) (Model II) was compared with a model with multi-CES and nonhomotheticity assumptions as specified in equation (18) (Model III). This is to jointly test whether or not a specification based on the multi-CES and nonhomotheticity 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 (Judge et al., 1982, pp. 326 – 328; 1985, pp. 472 – 477). For a detailed description of the test procedure, see Appendix A.

Data

Trade-flow data on rice were collected from the *Commodity Trade Statistics* (United Nations) for 25 years, from 1962 to 1986, based on calendar years. Imports by an individual country not only varied but were often zero in certain years, making econometric analysis more difficult. As a result, all the importing countries and regions were aggregated into one region. Exporters were categorized into seven regions: Thailand, the U.S.A., Argentina, Australia, Burma, Italy, and Pakistan.[10] Exporters that did not report their data to the United Nations were excluded.[11] U.S. government shipment data were collected from FATUS (U.S. Department of Agriculture). 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 were excluded, because actual prices for these types of shipments deviate considerably from market prices. Data for budgets allocated to imports are not available; therefore, V_i was approximated by total expenditure for imports of the products from individual suppliers. The consumer price index in the U.S. was employed as price index P^*, becaue V_i was expressed in U.S. dollars.

The prices used in this analysis are expressed in U.S. dollars and were calculated as the total value of exports divided by the total quantity exported from each exporting country. If the analysis were based on multiple importing regions, it would be preferable to use the landed prices of rice imported from the different exporting countries. Import prices would differ from the export prices used in this study by the amount of the total costs of transporting rice between the two regions. For this study, however, the various importing regions have been aggregated into one region so that specific bilateral transportation costs cannot be included. More importantly, the differences in transportation costs are reflected in the parameter γ_{ij}, which can be seen as an adjustment factor that transforms the single CES into country-specific multi-CES, σ_{ij}, as shown above. Accordingly, it may not be unreasonable to assume that the relevant price for analyzing a country's ability to export to the world as a whole is the price at which it exports the commodity.

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

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

It should be emphasized that data related to international trade in rice, as well as other commodities, are frequently unreliable. Moreover, data series that actually measure the variables of interest are often unavailable. These kinds of data problems mean that caution should be exercised in the interpretation of statistical results derived from quantitative models of international trade. Although some data problems were encountered in conducting this study, we feel that the statistical results are sufficiently robust to support the conclusions drawn.

Empirical Results

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. Serious correlations were observed when testing correlation coefficients among the error terms generated from ordinary least-square (OLS) equations for individual exporters. In addition, the estimated price coefficients were not the same but were statistically different from one another in the equations for individual exporters.[12] 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 Armington's single-CES assumption were compared. The quantity-dependent equation is the double-log form from equation (6) – (Model I); and the market share equation is the double-log model from equation (7) – (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 other. The R-squares in Model II generally are greater than those in Model I, while two of the R-squares in Model I are negative.[13]

Third, the appropriateness of the single-CES and homotheticity assumptions is tested. The results above indicate that a market-share-dependent specification appears to be superior to the quantity-dependent specification. Therefore, the system of equations in Model II, the original Armington

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

[13] The negative R-squares in Model I may be due to the restriction that price coefficients be identical for all suppliers, in addition to using quantity- (instead of market-share-) dependent specification using SUR.

TABLE 3

Comparison of different specifications among original and modified Armington procedures[a]

Method	Original Armington procedure				Modified Armington procedure		
Assumption	Single-CES and homotheticity				Multi-CES and nonhomotheticity		
Model dependent variable	Model I Quantitiy		Model II Market-share		Model III 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.A.	0.396	−1.683 (0.126)	0.553	−1.689 (0.111)	0.555	−1.519 (0.304)	0.082 (0.145)
Australia	0.315	−1.683 (0.126)	0.546	−1.689 (0.111)	0.576	−1.851 (0.340)	0.122 (0.185)
Burma	−0.067	−1.683 (0.126)	0.182	−1.689 (0.111)	0.444	−0.984 (0.489)	−1.291 (0.386)
Italy	−0.342	−1.683 (0.126)	0.058	−1.689 (0.111)	0.197	−1.284 (0.288)	0.388 (0.261)
Pakistan	0.473	−1.683 (0.126)	0.624	−1.689 (0.111)	0.724	−1.726 (0.301)	0.546 (0.229)

[a] The seemingly unrelated regression (SUR) is used in each model. Standard errors are in parentheses.

single-CES and homotheticity assumptions with market-share-dependent specification, was compared with that of Model III, multi-CES and nonhomotheticity assumptions based on equation (18). Both equations are specified with market-share-dependent variables. The F-test based on Appendix equation (A2) resulted in an estimated F-value equal to 3.451. This F-value is greater than 2.18, which is the critical value at the 1% significance level for degrees of freedom of $\nu_1 = 12$ and $\nu_2 = 121 \rightarrow \infty$ (Table 4). Accordingly, Armington's original specification with single-CES and homotheticity is statistically rejected at the 1% significance level. This indicates that the system of equations in Model III is statistically superior to

TABLE 4

Results of hypothesis testing for Armington's original assumptions

Σ_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

that of 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: Burma (negative coefficient) and Pakistan (positive coefficient). This implies that an assumption of $\beta_{ij} = 0$ for $\forall j$ is rejected and indicates that market shares of individual exporters are not always independent of changes in the level of budget expenditures allocated to the imports, 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.

The estimated coefficients in this study of world trade in rice 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 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, 1985). 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 does 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.

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 fact 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 imported rice.

Conclusions

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 were evaluated. The relevance of Armington's assumptions was examined, and a modified approach was 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 of this research are basically consistent with those found by Winters (1982) and Alston et al. (1989). However, these results do not necessarily imply that 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 instead of quantity as the dependent variable 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.

For this study, the first-stage model was not estimated. On the basis of the modified second-stage equations, some interesting empirical results were obtained. In particular, it appears that importers are quite sensitive to relative prices. This result is consistent with the fact that the world market for rice is quite small relative to total rice production (Barker et al., 1985). Importers in this thin market can choose among several current suppliers as well as potential exporters such as India and China. In choosing among alternative suppliers, the results of this analysis indicate that importers consider relative prices and the quality of the rice being imported. This observation is consistent with the results of studies by Ito et al. (1988). Overall, the approach suggested in this paper appears to be conceptually sound and useful in obtaining important empirical results.

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Appendix A

A test of superiority between two systems

To determine superiority of one system model to another, an F-test can be applied (Judge et al., 1982, pp. 326–328; 1985, pp. 472–477). In this research, two system models, with and without restrictions, are compared. The set of restrictions are expressed by:

$$\mathbf{RB} - \mathbf{r} = 0 \tag{A1}$$

where $\mathbf{R}$ and $\mathbf{r}$ are known matrices of dimensions $(J \times K)$ and $(J \times 1)$, respectively. The hypothesis is:

H_0: $\mathbf{RB} - \mathbf{r} = 0$
H_a: $\mathbf{RB} - \mathbf{r} \neq 0$

If H_0 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 (the single-CES estimate) and budged coefficients (the coefficients being equal to 0 under homotheticity) are imposed. On the other hand, the Model III system has no restrictions.

An F-test to determine superiority of one model to another with respect to restrictions is expressed as follows:

$$\lambda = A_J/B_{MT-K} \approx F_{J,\ MT-K} \tag{A2}$$

where

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

and

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

$\hat{\mathrm{B}}^*$ represents estimated copefficients under restrictions, and $\hat{\mathrm{B}}$ 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 restricitons, and T is the number of observations in each equation.

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