TESTING FOR AGGREGATION AND SIMULTANEOUS BIAS IN U.S. SOYBEAN EXPORT EQUATIONS

Carlos A. Arnade and Cecil W. Davison

Abstract

Most previous estimates of elasticities of export demand for U.S. soybeans have emanated from single import equations subject to aggregation and simultaneous equation bias. This analysis tests U.S. soybean export data for aggregation and simultaneous equation bias and divides the aggregate data into six market equations to reduce these biases. Elasticity estimates from the six equations are compared with elasticity estimates from single equation OLS and 2SLS estimations using the same aggregate data. Results suggest that distortions from unjustified 2SLS estimation may exceed those from aggregation bias.

Key words: aggregation bias, simultaneous equation bias, soybean exports, price elasticity, market share.

Policymakers, exporters, and researchers are interested in export elasticity estimates that most accurately reflect importers' responses to changes in important explanatory variables, particularly price. Previous estimates of the short-run price elasticity of demand for U.S. soybean exports, reviewed by Gardiner and Dixit, range from inelastic (-0.14) to elastic (-2.00) with no consensus on the appropriate range and are estimated from aggregate data (summed across countries), which could distort the estimates with bias from aggregation and simultaneity.

This analysis tests for both aggregation bias and simultaneous equation bias in import demand equations. The article then presents elasticity estimates compiled from specific markets, in order to reduce the effects of aggregation and simultaneous equation bias. Finally, it compares a weighted sum of market specific elasticity estimates with estimates from single equation ordinary least squares (OLS) and two-stage least squares (2SLS) estimations.

BACKGROUND

Typically, elasticity estimates vary because of differences in: estimation methods, model specification, the time period of estimation, the type of data (quarterly or annual), and the quality of data available to researchers. Variations in specification, time periods, and data are expected among published elasticity estimates, and can obscure variations due to methods employed and data aggregation.

Research summarized by Gardiner and Dixit used data aggregated across importing countries in one or a few equations (characterized as a single equation approach in this article) to obtain estimates of export demand elasticities. For example, Houck et al. used aggregate data in a single import demand equation and obtained elasticity estimates by OLS, 2SLS, and 3SLS estimators. Chambers and Just used aggregate data in single import demand equations as part of a 3SLS system of simultaneous equations. Aggregate data are subject to inherent problems that include the following:

(1) Simultaneous equation bias is likely when U.S. exports are aggregated. Imports of U.S. soybeans by one or two countries may not influence U.S. prices, but imports by all countries may.

(2) Aggregation bias will occur if the parameters on the linearly aggregated exogenous variables are not the same across individual demand equations (Zellner).

(3) A single equation requires a broad exchange rate index, whereas country-specific exchange rates can be used in individual market equations. Thus, market-

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Senior authorship is not assigned.

The authors thank Joe Glauber, Stephen Haley, Paul Johnston, Stephanie Mercier, Stephen Miller, and three anonymous *Journal* referees for helpful comments and suggestions.

specific equations avoid the generalities of broad-based indexes.

This article first presents the conditions for simple linear aggregation of demand equations, then tests for evidence of simultaneous equation bias (problem [1]) and aggregation bias (problem [2]), and finally presents a multiple-equation estimation procedure to reduce the effects of both types of bias. The market-specific multiple-equation estimation procedure presented herein provides estimates of specific exchange rate effects on U.S. soybean exports to individual major markets, thus addressing problem [3].

CONDITIONS FOR AGGREGATION

There are several ways of demonstrating the conditions for aggregation. Deaton and Muellbauer (pp. 148–53) demonstrate the conditions for aggregating individual consumer demand functions whose arguments are prices and total expenditures. They point out that linearly aggregated demand functions are subject to aggregation bias if aggregate demand is a function of the distribution of expenditures across consumers as well as the level of aggregate expenditures. We provide a simple demonstration of sufficient conditions required for aggregating *any* two demand functions whose arguments are prices and *any* other variables. These demand functions may represent the import demand of two different countries as well as being input or consumer demand functions.

Suppose the demand functions for two countries are linear in price and another variable such as income. The demand for country i is

$$(1)\ d_i = b_iP + a_iY_i,$$

where d_i is the quantity purchased in the *ith* country as a function of world price (P) and that country's income (Y_i). The demand for country j is

$$(2)\ d_j = b_jP + a_jY_j,$$

where d_j is the quantity purchased in the *jth* country as a function of price (P) and the *jth* country's income (Y_j). Aggregate demand (D), expressed as

$$(3)\ D = BP + AY,$$

is a function of price (P) and aggregate income (Y), and, by definition, equals the sum of the individual country demand functions,

$$(4)\ D = d_i + d_j.$$

Substituting terms from equations (1), (2), and (3) into equation (4) yields

$$(5)\ BP + AY = (b_iP + a_iY_i) + (b_jP + a_jY_j).$$

Assume that the price effects (BP) in the aggregate demand function, equation (3), equal the sum of the price effects in the individual demand functions,

$$(6)\ BP = b_iP + b_jP.$$

Subtracting equation (6) from (5) shows that the income effects in D equal the sum of the income effects in d_i and d_j,

$$(7)\ AY = a_iY_i + a_jY_j.$$

Dividing both sides of equation (6) by P simply shows that the sum of the parameters on P equals the aggregate parameter on price, or that

$$(8)\ B = b_i + b_j.$$

By definition, aggregate income equals the sum of income in the two countries,

$$(9)\ Y = Y_i + Y_j.$$

Substituting terms from equation (9) into (7) gives

$$(10)\ AY = A(Y_i + Y_j) = AY_i + AY_j = a_iY_i + a_jY_j,$$

which is true when $a_i = a_j$. Furthermore, Deaton and Muellbauer state that for exact linear aggregation, the parameters on the Y term *must* be equal in each equation (p. 150). Zellner affirms (without the simplifying assumption of equation (6)) that there will be no aggregation bias involved in simple linear aggregation if the parameters on income are equal across individual demand functions. However, this argument applies not only to income aggregated across individuals but to any variable summed across equations or individual countries.

Applying these conditions for linear aggregation of demand functions to linear aggregation of import demand functions, we derive the null hypothesis to test for evidence of aggregation bias: parameters on all the linearly aggregated exogenous variables are the same across market-specific import demand equations.

METHOD

Nineteen soybean importing countries used by Stallings in constructing his trade-weighted real exchange rate index for U.S. soybean markets were used to estimate six equations with annual data for U.S. soybean exports to the EC-9, Japan, Spain, Taiwan, South Korea, and the remainder of the 19 countries (rest of world, ROW), which collectively imported 93 percent of U.S. soybean exports during Stalling's 1983–1985 base period (countries and trade weights in Appendix).

Previous studies of export demand for U.S. soybeans have included as their explanatory variables the price of soybean meal as a substitute for soybeans (Houck and Mann; Houck et al.), income or livestock in the importing countries (Houck and Mann; Houck et al.; Helmberger and Akinyosoye; Chambers and Just), and exchange rates (Anderson; Chambers and Just). We specified our soybean import demand equations as input demand equations with the prices of U.S. soybeans and soybean meal, an exchange rate index, and pork production (as a measure of output) as explanatory variables.

We chose pork production as a representative of livestock production that uses soybean meal in foreign countries, excluding ruminant meat production that uses forages more extensively than high-protein concentrate rations in foreign countries. Poultry production, especially broiler production, also uses soybean meal in feed rations. However, the largest importer of U.S. soybeans, the EC, uses substantially more oilseed in pork production than in poultry meat production (Leuck).

The six equations were specified as linear combinations of the exogenous variables and estimated in the form

$$(11)\ SBX_i = b_0 + b_{1i}SBP + b_{2i}SMP + b_{3i}PORK_i + b_{4i}EXR_i + u_i,$$

where

SBX_i = U.S. soybean exports to the *ith* market (i = 1,...6);

SBP = U.S. soybean price, Rotterdam ($/metric ton * 1/U.S. CPI);

SMP = U.S. soybean meal price, Rotterdam (ditto);

$PORK_i$ = pork production in the *ith* market;

EXR_i = real exchange rate index for the *ith* market: (foreign currency units/foreign CPI)/($1/U.S. CPI) indexed to 1980 = 100. For the EC and the ROW, the individual country's real exchange rate indexes were weighted by Stalling's trade weights (shares of U.S. soybean exports) before summing to aggregate indexes for the EC and the ROW (countries and weights in Appendix). The exchange rate indexes in these six markets, when weighted by Stalling's market shares, sum to Stalling's trade weighted exchange rate index, used in the OLS and 2SLS single equations for all 19 markets;

b_{ji} = parameters; and

u_i = normally distributed random errors.

Calendar year U.S. soybean exports, 1963–1986, were the dependent variables (United Nations, *Commodity Trade Statistics*). Soybean and soybean meal prices, exchange rates, and CPI indexes came from the International Monetary Fund's *International Financial Statistics* and Taiwan's statistical counterparts (Central Bank of China; Council for Economic Planning and Development). Pork (pigmeat) production came from computer tapes from the Food and Agriculture Organization of the United Nations. Zellner's unrestricted seemingly unrelated regression (SUR), using annual data, provided individual estimates of the parameters on the variables for all six equations.

Testing for Simultaneous Equation Bias

Before estimating our equations by SUR, we tested the market equation that represented the largest share of 1983–85 U.S. soybean exports (the EC, which averaged 36 percent) for simultaneous equation bias between soybean and soybean meal prices and U.S. soybean exports using a test developed by Wu and described by Chow (p. 314).

To test whether U.S. soybean and soybean meal prices were exogenous to EC imports of U.S. soybeans, we obtained instrumental variables for soybean and soybean meal prices whose estimated values were specified as a function of U.S. soybean exports to the EC, plus the additional explanatory variable of the price in t - 1.

We used the instrumental variables as exogenous variables in the EC import demand equation and obtained 2SLS estimates for the EC equation. We also obtained OLS estimates of the EC equation. Wu's statistic for testing for differences between 2SLS and OLS estimates in econometric equations,

$(H_0: B_{2S} = B_{OLS}$ against $H_a: B_{2S} \neq B_{OLS})$,

is

$$W = n(B_{2S} - B_{OLS})'V(q)^{-1}(B_{2S} - B_{OLS}),$$

where

n = number of observations;

B_{2S} = a vector of the 2SLS estimates of interest;

B_{OLS} = a vector of OLS estimates of interest; and

$V(q)$ = the variance-covariance matrix of the vector $(B_{2S} - B_{OLS})$, represented by $n(V_{2S} - V_{OLS})$ or n times the differences between variance-covariance matrices of 2SLS and OLS estimates.

The statistic W has a χ^2 with one degree of freedom as its asymptotic distribution if the null hypothesis is true.

The Wu test produced no evidence of differences between the 2SLS and OLS estimates of the EC equation. (The W statistic calculated for the soybean and soybean meal prices in the EC equation was 1.31, which is not significant at the five-percent level.) We concluded that EC imports did not influence the U.S. soybean and soybean meal prices and assumed the prices are exogenous. We then assumed that the other five markets, whose shares ranged from 25 percent down to 4 percent, were also price takers and that the U.S. soybean and soybean meal prices were exogenous to all six equations.

We applied the same test to a market-wide equation using data aggregated across all 19 countries. The W statistic calculated for soybean and soybean meal prices in the equation was 0.67, which is not significant at the five-percent level. Thus, the W statistics calculated for both the EC and world (19-country) equations were insufficient at the five-percent level to suggest that the coefficients are subject to simultaneous equation bias. Consequently, 2SLS estimates are not appropriate. However, 2SLS estimation of the world equation is a commonly accepted, if not recommended, procedure (in the absence of the Wu test). Therefore, we present 2SLS estimation results to compare with the OLS and SUR estimation results.

Testing for Aggregation Bias

To test the null hypothesis that the parameters on the linearly aggregated variable (pork production) were the same across all six equations,

$$H_0: b_{31} = b_{32} = ... = b_{36},$$

versus the alternative,

$$H_a: \text{at least one } b_{3i} \neq b_{3j}, (i \neq j),$$

where b_{3i} is the parameter on pork production in the *ith* equation, the six market specific equations were estimated first by SUR without any restrictions. Then the equations were reestimated with the restriction that the estimators on the pork production variable were the same across all six equations.

Testing the results of this restriction determines whether we consider the parameters on the aggregated variable the same across the individual markets. If the restriction on the pork production estimator significantly alters the variance-covariance matrix of errors between the six equations, we can reject H_0 and conclude that estimates from a single equation would contain aggregation bias.

To determine if the restricted estimations were significantly different from the unrestricted, the statistic g was used,

$$g = (r - R\hat{B})'(RCR')^{-1}(r - R\hat{B}),$$

where $r = RB$ represents a matrix of linear restrictions on the coefficient vector B; $C = [X'(\Sigma^{-1} \otimes I)X]^{-1}$; X represents the matrix of the exogenous variables; Σ is the variance-covariance matrix of errors between equations; I is an identity matrix; and $\otimes$ denotes Kronecker product (Judge et al., p. 28). The statistic g is χ^2 distributed, with degrees of freedom equal to the number of restrictions (five in our case). In deriving and estimating g, Σ^{-1} is replaced by $\hat{\Sigma}^{-1}$ (see Judge et al., pp. 472–76). Our calculated g statistic was 62, significant at the one percent level, leading to rejection of the hypothesis that the parameters on all the aggregated variables are the same across country-specific markets. Thus, one of the conditions for using aggregate data to estimate a single equation is violated.

Weighted-Market-Share Estimation

From our aggregation-bias test, we concluded that single equation estimation using this export demand data, aggregated across country-specific markets, contains aggregation bias. Comparing single-equation elasticity estimates with trade-weighted elasticities from the six equations may help reveal the extent of the bias.

Parameter estimates and t-values from the unrestricted SUR estimations of the six market-specific equations are in Table 1. Results from a total-export single OLS equation and a total-export 2SLS system of equations are in Table 2. Elasticities, calculated at the sample means from each of the six SUR equations, were first weighted by that market's share of U.S. soybean exports for 1983–85, and then added to obtain aggregate U.S. elasticity estimates across all

TABLE 1. SUR ESTIMATION RESULTS FOR U.S. SOYBEAN EXPORT EQUATIONS, 1963-86[a]

Variables/data	EC-9	Japan	Spain	Taiwan	South Korea	Rest of World[b]
Constant	2,391,000	435,600	1,676,000	751,000	38,670	-2,479,000
	(1.22)	(.58)	(2.96)**	(2.10)*	(.49)	(-3.17)**
Real U.S. soybean price, Rotterdam[c]	-5,469	-3,045	-1,918	-877.6	-1,462	2,452
	(-.98)	(-2.42)*	(-.95)	(-1.48)	(-2.52)*	(1.23)
Real U.S. soybean meal price, Rotterdam[c]	4,258	3,522	287.3	304.5	889.3	-2,659
	(.89)	(3.02)**	(.17)	(.52)	(1.95)*	(-1.56)
Pork production in importing country[d]	1.042	2.338	.786	1.807	1.967	.949
	(5.52)**	(7.37)**	(3.35)**	(13.73)**	(7.99)**	(9.03)**
Real exchange rate index	-37,080	981.6	-8,710	-4,180	564.2	23,290
	(-3.66)**	(.30)	(-2.71)**	(-3.04)**	(.56)	(4.29)
$\bar{R}^2$	.71	.92	.54	.95	.76	.78
Durbin-Watson	2.01	2.00	1.71	2.22	1.76	1.84
F[e]	366**	684**	341**	759**	348**	354**

[a] t-values in parentheses. Significance levels (one-tailed test with 19 degrees of freedom, two-tailed test for constant): * = 5 percent, ** = 1 percent.
[b] Mexico, Portugal, Israel, Switzerland, Canada, Norway, Greece, Indonesia, and Egypt.
[c] Jan.-Dec. average, dollars per metric ton, deflated by U.S. CPI.
[d] Metric tons.
[e] Test of significance of model (unrestricted SUR estimation compared with restriction of estimators = 0 in each equation in turn).

TABLE 2. OLS AND 2SLS ESTIMATION RESULTS[a]

Variables/data	Estimator: OLS	Estimator: 2SLS system		
Dependent variable	U.S. soybean exports to world[b]	Real U.S. soybean price	Real U.S. soybean meal price	U.S. soybean exports to world[b]
Constant	6,324,000	111.4	155.6	9,609,000
	(1.44)	(1.78)	(2.50)*	(1.54)
U.S. soybean exports to world[b]	—	-.0000019	-.0000016	—
		(-.76)	(-.56)	
Real U.S. soybean price, Rotterdam[d]	-13,950	.6846	—	-4,127[c]
	(-1.51)	(4.03)**		(-.56)
Real U.S. soybean meal price, Rotterdam[d]	7,015	—	.4308	-23,360[c]
	(.92)		(2.17)*	(-1.17)
Pork production in importing country[e]	1.193	—	—	1.157
	(5.82)**			(5.04)**
Real exchange rate index	-64,630	—	—	-60,430
	(-2.60)**			(-2.19)*
R^2	.87	.39	.11	—
Durbin-Watson	2.06	—	—	—
F[f]	30**	5.82**	1.67	—
Degrees of freedom	19	22	22	19

[a] t-values in parentheses. Significance levels (two-tailed test on constants and instrument equations, one-tailed test on other variables): * = 5 percent, ** = 1 percent. In the price (instrument) equations, the dependent variable is lagged 1 year on the right-hand side.
[b] Nineteen countries listed in Appendix.
[c] Estimate of parameter on instrumental variable.
[d] Jan.-Dec. average, dollars per metric ton, deflated by U.S. CPI.
[e] Metric tons.
[f] Test of significance of model [R^2/no. of variables]/[(1 - R^2)/(no. of observations - no. of variables)].

TABLE 3. PRICE, CROSS-PRICE, PORK, AND EXCHANGE RATE INDEX ELASTICITIES FOR U.S. SOYBEAN EXPORTS

Market	Soybean price	Soybean meal price	Pork	Exchange rate index	Market share [a]	Soybean price	Soybean meal price	Pork	Exchange rate index
	---------- Elasticities [b] ----------					------- Weighted elasticities [c] -------			
SUR Estimation									
EC-9	-0.288	0.183	1.397	-0.930	0.358	-0.103	0.066	0.500	-0.333
Japan	-.303	.287	.797	.046 [d]	.246	-.075	.071	.196	.011
Spain	-.475	.058	.505	-1.105	.086	-.041	.005	.043	-.095
Taiwan	-.351	.100	1.196	-.698	.079	-.028	.008	.094	-.055
S. Korea	-1.720	.857	1.367	.256 [d]	.043	-.074	.037	.059	.011
Rest of World [e]	.375 [d]	-.333 [d]	1.039	1.452 [d]	.187	.070	-.062	.194	.272
World Total [f]					.999				
Plausible Sign						-.32	.19	1.09	-.48
All						-.25	.12	1.09	-.19
Ordinary Least Squares Estimation						---------- Elasticities [b] ----------			
						-.32	.13	1.14	-.68
Two-stage Least Squares Estimation						-.10	-.45	1.10	-.64

[a] Average share of the U.S. export market, 1983–85 (Appendix).
[b] Calculated at the sample means (Appendix).
[c] Elasticities times market share, computed from unrounded data.
[d] Implausible sign.
[e] Mexico, Portugal, Israel, Switzerland, Canada, Norway, Greece, Indonesia, and Egypt.
[f] Nineteen countries that imported most of U.S. soybeans (Appendix).

markets (Table 3). The elasticities were summed two ways: first by totaling all that had the expected sign, and then by including the implausibly-signed estimates, which changed the price, cross-price, and exchange rate index elasticity estimates by 22, 37, and 60 percent, respectively.

Elasticity estimates from the six-equation estimation may contain elements of aggregation bias from the EC and ROW equations. Within the six country markets aggregated for the EC equation, one might expect similarly sloped expansion paths at various levels of pork production because these EC countries are geographic and economic neighbors and have similar standards of living. Consequently, one may not expect serious aggregation bias effects in the EC equation. The ROW equation, however, contains nine diverse countries which span continents and range from developed to developing economies (countries in Appendix). Expansion paths at various levels of pork production could not be expected to be as similar across the ROW countries as in the EC. Hence, one would expect greater effects of aggregation bias in the ROW equation than in the EC equation. However, in the six-equation weighted-market-share elasticity estimation (Table 3), the ROW market share is only 19 percent, and the ROW elasticities are weighted accordingly.

RESULTS

Price and cross-price elasticity estimates from the SUR six-equation estimation are closer to those from the OLS single-equation estimation (which probably contains aggregation bias but did not reveal evidence of simultaneous equation bias) than to the 2SLS estimates (Table 3). The 2SLS estimation, normally used to correct for simultaneous equation bias (assumed or otherwise), also probably contains aggregation bias. If total exports influence the U.S. price, conventional econometric procedures would suggest the 2SLS estimates are better than the OLS estimates. However, in this case, the 2SLS estimation appears to introduce distortions in the price and cross-price elasticity estimates that exceed those that may be attributed to aggregation bias.

Our OLS and six-equation deflated soybean price elasticity estimates of -0.32 and -0.25 are lower than estimates by Houck et al. (p. 86, OLS = -0.53, 2SLS = -0.54, 3SLS = -0.67, -0.68), who used annual data for 1946–1966 (price variable = soybean price/soybean meal price). Our 2SLS estimate of -0.10 is lower than Chambers and Just's 3SLS estimate of -0.20 (quarterly data, 1969:1-1977:2, deflated prices), close to Helmberger and Akinyosoye's 3SLS estimate of -0.14 (annual data, 1948/49–1977/78, deflated prices), but lower than Houck and Mann's 2SLS estimate of -0.32 (annual data, 1946–1964, nominal prices). Conway, using a stochastic coefficients approach to reestimate Chambers and Just's quarterly model (omitting the seasonal variables), confirmed their estimated soybean price elasticity of -0.20. All of these other published estimates were from single-equation estimations, which were subject to aggregation bias, as are our OLS and 2SLS estimates.

Our -0.30 deflated soybean price elasticity estimate for U.S. exports to Japan is lower than Greenshields' -0.65 (annual data, 1955–73, deflated import price index), but close to the -0.35 estimate by Meyers et al. (annual data, 1960/61–1976/77; elasticities for 1973/74–1976/77, price variable = soybean wholesale price index in Japan).

Our soybean price elasticity estimate of -0.29 for the EC exceeds the -0.23 estimate by Knipscheer et al. (semi-annual data, 1961–1976, price variable = soybean meal price/corn price). We would expect our elasticity estimate to exceed theirs because their dependent variable was total EC imports of both soybeans and soybean meal (per animal feed unit), the demand for which would be less elastic than for total soybeans alone, which would be less elastic than the EC demand for U.S. soybeans. (U.S. soybeans constituted 77 percent of EC soybean imports, 1974–1985 [Davison]). Also, we would expect a one-year elasticity to exceed a six-month elasticity.

CONCLUSIONS

Estimating export demand for U.S. soybeans in a single equation using data aggregated across all markets subjects the estimates to both aggregation and simultaneous equation bias. The prevalence of import equations estimated by 2SLS or 3SLS in the literature indicates awareness of and correction for simultaneous equation bias. However, across-country aggregation bias seems to have attracted less attention.

If the aggregated variables and their parameters, plus the other exogenous variables, are the same across the individual markets in the correct specification, single-equation estimation is the quickest and easiest way of estimating the elasticities. If the parameters on the aggregated variables are not the same across the markets, as this study suggests, then aggregating individual-market data to estimate a single OLS or 2SLS import equation imposes unrealistic assumptions that may distort the estimates of the true elasticities.

Testing for evidence of simultaneous equation bias before accepting 2SLS estimates could obviate 2SLS distortions, which in this example appear to exceed those from aggregation bias.

The multiple-equation weighted-market-share approach, which reduces the problems of aggregation and simultaneous equation bias intrinsic to a single equation, requires more data but has the advantage of providing market-specific elasticity estimates that can be evaluated individually. Questionable equations or estimates can be identified and isolated. Researchers can then reestimate weak equations or use market-specific elasticities judged more appropriate.

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APPENDIX. U.S. SOYBEAN EXPORT-SHARE WEIGHTS USED IN TRADE-WEIGHTED REAL EXCHANGE RATE INDEXES FOR WORLD, EC-9, AND ROW EQUATIONS; PLUS SAMPLE MEANS

Country	Share-weight [a]	Sample means: U.S. soybean Price [b]	Sample means: U.S. soybean Meal price [b]	Sample means: U.S. soybean Exports [c]	Sample means: Pork production [d]	Sample means: Exchange rate index [e]
		Dollars/metric ton		Metric tons		1980=100
EC-9	0.358	293	240	5,575,228	7,475,988	140
Netherlands	.180					
W. Germany	.057					
Belgium-Lux	.044					
Italy	.033					
France	.023					
UK	.021					
Denmark	0					
Ireland	0					
Japan	.246	293	240	2,946,827	1,005,012	138
Spain	.086	293	240	1,185,031	761,986	150
Taiwan	.079	293	240	732,628	484,932	122
S. Korea	.043	293	240	249,282	173,194	113
Rest of World	.187	293	240	1,917,510	2,098,746	120
Mexico	.068					
Portugal	.033					
Israel	.023					
Switzerland	.014					
Canada	.013					
Norway	.013					
Greece	.011					
Indonesia	.011					
Egypt	.001					
Total	.999	293	240	12,606,505	11,999,857	133

[a] 1983-85 share of U.S. soybean exports, from Stallings.
[b] Deflated by U.S. CPI, from International Monetary Fund's *International Financial Statistics* (IFS).
[c] From United Nation's *Commodity Trade Statistics*.
[d] From FAO *Production Yearbooks*.
[e] Exchange rates deflated by U.S. and foreign CPI's (from *IFS*), indexed to 1980 = 100.