Large Versus Small Price Changes and the Demand for Imports

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THE PURPOSE OF THIS PAPER is twofold: first, to provide new estimates of the aggregate import demand equation for 12 industrial countries, using quarterly data on the relevant variables for the period 1955 to 1973; and, second, to examine empirically whether the elasticity of imports with respect to relative prices,[1] and the speed at which actual imports adjust to the desired level, are both *independent* of the size of the relative price change.

Previous studies that have estimated aggregate import functions for a large number of industrial countries have generally been restricted to using either annual data (e.g., see Heien (1968), Houthakker and Magee (1969), and Taplin (1972)) or semiannual data (e.g., see Samuelson (1973) and Khan and Ross (1975)).[2] Apart from the increased degrees of freedom that the use of quarterly data provides,[3] it should be possible with such data to make more precise estimates of the timing of the relationships involved.[4]

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[1] That is, the ratio of import prices to domestic prices.

[2] An exception is the study by Adams and others (1969), which estimated quarterly equations for total imports of 9 industrial countries for the period 1955–65.

[3] Although quarterly data on import values and import prices exist for all 12 countries in our sample, quarterly data on real or nominal income do not for most countries; therefore, it was necessary to interpolate the annual data to obtain quarterly estimates for the real income and domestic price variables. (See Appendix II.)

[4] See Magee (1975).

A second characteristic of most previous econometric studies of the demand for imports is the implicit assumption that the elasticities of import demand with respect to relative prices and real domestic income are *constant* for all values of the two explanatory variables. This assumption is, of course, inherent in the use of a log-linear functional form for the import equation. Since economic theory offers little guidance on the appropriate functional form for the import demand relationship, and since the double-log specification (in contrast to a linear specification) yields direct estimates of the relative price and real income elasticities for imports, it is a convenient form for purposes of estimation.[5]

Beginning, however, with Orcutt's classic paper (1950) on the measurement of price elasticities in international trade, several writers have argued that the relative price elasticity of demand for imports will be different (larger) for large price changes than for small price changes. The argument has appeared in two alternative forms. One interpretation is that the relative price elasticity itself will vary directly with the size of the price change, since the price change must be large enough to overcome buyer inertia and the costs associated with switching suppliers (e.g., see Orcutt (1950) and Kindleberger (1973)).[6] A representative statement of this position has been given by Liu (1954, p. 437):

> . . . There are always "costs" involved in a change of the things a business firm has become accustomed to purchase. New contacts must be made; the quality of the new product tested . . . A one-point change in relative prices may not justify the change-over at all. A five-point change may make a partial or complete switch-over unavoidable for some importers. A ten-point change would not only increase the extent of substitution for those who had made a partial change, but also increase the number who must for the first time make a shift.

The second interpretation is that the adjustment of import quantity to large price changes is more rapid than the adjustment to small changes.[7] Under this interpretation, the long-run (or equilibrium) elasticities of relative prices and real income would be independent of the size of the price change, but the short-run (or impact) elasticities would be functionally related to the price change itself.[8] Stated in other words, the

[5] For an empirical examination of the appropriate functional form, see Khan and Ross (1974).

[6] Some writers—for example, Leamer and Stern (1970)—have objected to this argument on the grounds that it is inconsistent with the basic assumption of a long-run demand function that depends only on current prices and not on the history of prices.

[7] This seems to be the interpretation favored, for example, by Leamer and Stern (1970), p. 34.

[8] If the speed of adjustment varies directly with the size of the price change, estimation of the standard, partial-adjustment import demand model will result in underestimation of the speed of adjustment but will not affect the total adjust-

response lag of importers to changes in the explanatory variables will vary inversely with the extent of the relative price change.

Despite the obvious importance of the issue of large versus small price changes, say, for forecasting the impact of exchange rate changes on trade flows,[9] little empirical work has been done on this question. In fact, the only published empirical study directly concerned with this hypothesis for aggregate import demand is apparently the study by Liu (1954) for U.S. imports. In brief, Liu included a squared relative price change variable in the import demand function (in addition to the usual income and relative price variables) to test for a nonlinearity in the relative price elasticity of import demand. Liu found the coefficient of the squared price change variable to be significantly different from zero; in fact, it was much better determined than was the estimated coefficient of the usual relative price variable. He therefore concluded that the large/small price change distinction was valid and that relative price changes had to be of a certain minimum size before they affected the demand for imports. Liu's results, however, should be regarded with some caution, because his estimates were restricted to the United States for a limited number of observations (17), and also because he made no distinction between the price-elasticity and adjustment speed versions of the large/small hypothesis.

The plan of this paper is as follows. Section I presents the four alternative specifications of the aggregate import demand function that are to be estimated. Two of these specifications are identical to those commonly found in the econometric literature on import demand: both the real income and relative price elasticities, and the speed of adjustment are assumed to be constant. The other two specifications allow the relative price elasticity and the speed of adjustment to vary with the change in relative prices. Section II presents the estimated results for these four import models and then compares the estimated price and income elasticities, both short-run and long-run, with earlier estimates for the same 12 countries: Belgium, Denmark, France, Finland, the Federal Republic of Germany, Italy, Japan, the Netherlands, Norway,

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ment. That is, the long-run income and relative-price elasticities will still be properly estimated, but the short-run (impact) elasticities will not; see Leamer and Stern (1970), p. 34.

[9] Changes in the exchange rate would generally result in larger changes in relative prices than would normally occur. In addition, the relative price elasticity associated with exchange rate changes might differ from that for other price changes (even of the same size) if exchange-rate induced changes are viewed as more *permanent* than other changes. The recent empirical study by Junz and Rhomberg (1973), however, found no empirical support for the hypothesis that the response of trade flows (export market shares) to exchange-rate induced price changes was different from the response to price changes measured in local currency.

Sweden, the United Kingdom, and the United States. The period of study is from 1955 to 1973 on a quarterly basis. In an effort to gauge the robustness of the original findings, Section III reports the results of some further empirical tests of the large/small hypothesis. Section IV contains a summary of the study's results as well as the principal conclusions. Appendix I contains three tables that provide additional estimates of the import demand function using either more disaggregated import data or an alternative time period. Appendix II gives the definitions of the variables and the data sources used in the equations.

I. The Aggregate Import Demand Function

The simplest import demand function relates the quantity of imports demanded by a country to the ratio of import prices to domestic prices,[10] and to the level of real income in that country.[11] In terms of logarithms, the equation can be specified as follows:

$$\log M_t^d = a_0 + a_1 \log P_t + a_2 \log Y_t + u_t \tag{1}$$

where

M = quantity of imports in period t,

P = ratio of import prices (PM) to domestic prices (PD) in period t, that is, $P_t = \frac{PM}{PD}\, t$,

Y = evel of gross national product (GNP) in constant prices in period t,

u is an error term, and the superscript d signifies demand.

Since equation (1) is specified in terms of logarithms, a_1 and a_2 represent the relative price and the real income elasticities of imports, respectively. It is expected that the price elasticity, a_1, will have a negative sign, but the sign of the income elasticity is ambiguous. Generally the sign is taken to be positive (as real income rises, the quantity of imports demanded will increase), but there is a possibility that the sign could be negative. For example, if imports are viewed as representing

[10] Assuming that imports are a substitute for domestic goods.

[11] See Houthakker and Magee (1969) and Leamer and Stern (1970). There is some controversy as to whether real income or real expenditure should be used as the relevant explanatory variable. The former excludes imports but includes exports, which represent the rest of the world's demand for the country's goods. Real expenditure includes imports and excludes exports. Most writers have chosen to work with real income, and we will follow this precedent.

the difference between the consumption and the production of importables, and if production rises faster in response to a rise in real income than does consumption, imports could easily fall as real income rises and thus yield a negative sign for a_2.[12]

One reason why the import demand equation is usually specified in logarithmic form is that this form allows imports to react proportionally to a rise and fall in the explanatory variables; that is, on the assumption of constant elasticities, the logarithmic form avoids the problem of changes in the elasticities as import quantities change.[13]

Since estimating equation (1) directly would imply that there is no time lag in the adjustment of actual imports (M_t) to import demand ($M^d{}_t$), we next relax this restrictive assumption by specifying a partial-adjustment mechanism for imports. This mechanism relates the change in imports in period t to the difference between the demand for imports in that period and the actual flow of imports in the previous period, $t - 1$:

$$\Delta \log M_t = \gamma[\log M_t^d - \log M_{t-1}] \tag{2}$$

where $\Delta \log M_t = \log M_t - \log M_{t-1}$, and γ is the coefficient of adjustment:

$$0 \leq \gamma \leq 1$$

When applied to the imports of a particular country, this partial-adjustment mechanism implicitly assumes that import prices are exogenous to the home country (usually being determined abroad) with quantities being adjusted domestically. A theoretical rationale for equation (2) is generally made on the grounds that costs are involved in the adjustment of actual imports to the desired flow, and that these costs constrain instantaneous adjustment.[14] Further, some imports may be linked to contracts extending over a period of time and thus cannot respond promptly to changes in demand.[15] This partial-adjustment framework introduces a distributed-lag structure with geometrically declining weights into the determination of imports, and thus the formulation is able to capture delayed response.[16]

[12] For a theoretical discussion of this, see Magee (1973). An empirical examination of this issue has been made by Khan and Ross (1975).

[13] For a discussion, see Khan and Ross (1974).

[14] These costs would generally be "search" costs or transactions costs; see Griliches (1967). For applications to imports, see Turnovsky (1968) and Khan (1974).

[15] See Malinvaud (1969).

[16] Equation (2) can be written in lag-operator notation as follows:

$$M_t = \frac{\gamma}{1 - wL} M_t^d \tag{2'}$$

(Cont'd. on p. 205.)

Substituting equation (1) in equation (2) and solving for imports in period t, we obtain:

$$\log M_t = \gamma a_0 + \gamma a_1 \log P_t + \gamma a_2 \log Y_t + (1 - \gamma) \log M_{t-1} + \gamma u_t \tag{3}$$

where γa_1 and γa_2 are taken to be the short-run, or "impact," price and income elasticities, respectively.

There would be several sources of potential bias in the estimates of the parameters if one were to estimate equations (1) and (3) directly by ordinary least squares. For example, a bias could result from simultaneity between imports and one or both of the explanatory variables. Biased estimates could also result from problems of aggregation as well as from errors in measurement of the relevant variables. These sources of bias have, however, already been dealt with rather extensively in the empirical trade literature since the publication of Orcutt's 1950 article.[17] One source of bias, also originally raised by Orcutt, that has not received much attention in the literature is the large/small price issue discussed earlier, that is, the possibility that the (relative) price elasticity of demand for imports is not constant but rather varies with the size of the price change. This is what Magee (1975) has termed the "quantum effect."

In this study, attention is concentrated on the two interpretations of the quantum effect that have already been described briefly. The first interpretation is the one directly attributable to Orcutt (1950), namely, that the price elasticity in equation (1), a_1, would be a function of the change in the relative price:

$$a_{1t} = \alpha_0 + \alpha_1 |\Delta \log P_t| + v_t \tag{4}$$

where Δ is a first-difference operator, $|\Delta \log P_t|$ is the absolute value of the change in relative price, and v is an error term. Since a larger price change would increase the price elasticity, it is expected that the coefficient α_1 would be negative.[18] If there is no influence of the change in relative prices on the price elasticity, we would expect that the constant term in equation (4) would be equal to the price elasticity:

$$\alpha_0 = a_1$$

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where $w = 1 - \gamma$ and L is a lag operator, $LM_t = M_{t-1}$. Equation (2′) is a representation of a distributed-lag structure with geometrically declining weights.

[17] The misspecification bias caused by ignoring possible time lags has been taken into account by our formulation of equation (3). For a discussion of the other sources of bias raised by Orcutt (1950), see Leamer and Stern (1970), Magee (1975), and Khan (1975).

[18] Our specification of equation (4) assumes that the relative price elasticity is independent of the direction (positive or negative) of the relative price change.

Substituting equation (4) in equation (1), we obtain:

$$\log M_t = a_0 + \alpha_0 \log P_t + \alpha_1 [|\Delta \log P_t| \cdot \log P_t] + a_2 \log Y_t + \eta_t \quad (5)$$

where $\eta_t = \log P_t \cdot v_t + u_t$

Assume that the error term in equation (4) is distributed with zero mean and variance, $\sigma_v^2\, \Omega$, independently of u_t. With these assumptions, the properties of η_t will be as follows:

$$E(\eta_t) = 0$$

$$Var(\eta_t) = \sigma_u^2 + \sigma_v^2 \cdot \log P_t \cdot \Omega$$

$$Cov(\eta_t \eta_s) = 0, \text{ for } t \neq s$$

When the variance of v_t is zero, the variance of η_t will be σ_u^a and equation (5) can be estimated by ordinary least squares. However, when the variance of v_t is nonzero, the η_t will be heteroscedastic. if Ω is known but the σ_u^2 and σ_v^2 are not, a generalization of Theil's two-step ordinary least-squares/generalized least-squares procedure can be used, although it involves a great deal of computation.[19] When Ω is not known, there seems to be no solution in the literature. Belsley (1973) has therefore suggested that for most economic problems it may be justifiable to assume that $\sigma_v^2\Omega = 0$. Our approach was to estimate equation (5) by ordinary least-squares (OLS) methods and to test for the presence of heteroscedasticity using the method attributable to Goldfeld and Quandt.[20] The OLS estimates will be unbiased, and the significance of α_1 can be considered a test of the quantum hypothesis. The hypothesis is rejected if $\alpha_1 = 0$, since in such a case equation (5) will simply reduce to equation (1).

The second interpretation of the quantum effect involves testing whether the coefficient of adjustment in equation (3) is a function of the size of the price change. It is assumed that importers adjust faster when larger price changes take place, and thus the mean time lag of adjustment is shortened.[21] We can therefore specify an equation of the form:

$$\gamma_t = \beta_0 + \beta_1 |\Delta \log P_t| + e_t \quad (6)$$

where $\beta_1 > 0$ and e is an error term.

Substituting equation (6) in equation (3), we obtain:

[19] See Theil (1971).
[20] See Goldfeld and Quandt (1970)
[21] The mean time lag is defined as $1/\gamma$.

$$\begin{aligned}\log M_t = {} & \beta_0 a_0 + \beta_0 a_1 \log P_t + \beta_0 a_2 \log Y_t + \\ & (1 - \beta_0) \log M_{t-1} + \beta_1 a_0 \, |\Delta \log P_t| + \\ & \beta_1 a_1 \, [|\Delta \log P_t| \cdot \log P_t] + \beta_1 a_2 \, [|\Delta \log P_t| \cdot \log Y_t] - \\ & \beta_1 \, [|\Delta \log P_t| \cdot \log M_{t-1}] + \epsilon_t \end{aligned} \tag{7}$$

$$\text{where } \epsilon_t = (a_0 + a_1 \log P_t + a_2 \log Y_t - \log M_{t-1}) e_t + (\beta_0 + \beta_1 \, |\Delta \log P_t| + e_t) u_t$$

Similar to the method described in discussing equation (5), we will estimate equation (7) by OLS and test the error term ϵ_t for heteroscedasticity. If $\beta_1 = 0$, equation (7) will be identical to equation (3).

In general, aggregate import equations have had serious problems with autocorrelation in the residuals,[22] a problem that would yield inefficient estimates of the parameters for equations (1) and (5), and both inefficient and biased estimates for equations (3) and (7)—owing to the presence of a lagged dependent variable.

To account for these possibilities of inefficiency and bias owing to autocorrelation, we estimated the four equations subject to the assumption that the errors in each follow a first-order autoregressive process of the form:

$$u_t = \rho_1 u_{t-1} + u_t^n \tag{8}$$

$$\gamma v_t = \rho_2 \gamma y_{t-1} + v_t^n \tag{9}$$

$$\eta_t = \rho_3 \eta_{t-1} + \eta_t^n \tag{10}$$

$$\epsilon_t = \rho_4 \epsilon_{t-1} + \epsilon_t^n \tag{11}$$

The ρ_i ($i = 1,2,3,4$) are the coefficients of autocorrelation, $|\rho_i| < 1$, and u_t^n and v_t^n are error terms with classical properties, that is, they are normal, independent, with zero means and constant variances. The η_t^n and ϵ_t^n also are normal, independent, with zero means, and, on the null hypothesis, constant variances.

II. Results: Total Imports, 1955–73

All four equations—(1), (3), (5), and (7)—were estimated by ordinary least squares using the two-step search method ascribed to Dhrymes (1971). This method provides efficient estimates of the parameters when the errors in the relationship follow a first-order autoregres-

[22] See Adams and others (1969), Houthakker and Magee (1969), and Khan and Ross (1975).

sive process.[23] The period of estimation was the third quarter of 1955 through the fourth quarter of 1973, and the data are seasonally adjusted. The precise definitions of the data and their sources are presented in Appendix II.

In discussing the results, we refer to the estimates of equations (1) and (5) as "equilibrium" estimates, since we assume there that demanded and actual imports are always equal. For obvious reasons, the results for equations (3) and (7) are termed the "disequilibrium" estimates. The results from estimating the equations are shown in Tables 1, 2, 4, and 5. Each table presents the estimates of the parameters, with their respective "*t*-values" in parentheses; the estimates of the coefficient of autocorrelation, ρ_i $(i = 1,2,3,4)$; the corrected coefficient of determination, $\bar{R}^2$; and the Durbin-Watson statistic, *D-W*.[24] We performed the likelihood test for heteroscedasticity for equations (5) and (7) and accepted the null hypothesis that the variance of the errors in each was constant. It was thus not necessary to adjust for heteroscedasticity in these equations.

The results for equation (1) are shown in Table 1. The estimated price elasticities have the expected negative signs and are significantly different from zero at the 5 per cent level in the equations for Belgium, Denmark, France, Finland, the Federal Republic of Germany, Norway, Sweden, and the United States. In the remaining four countries (Italy, the Netherlands, Japan, and the United Kingdom), the estimated price elasticity was not significantly different from zero (at the 5 per cent level) and for the latter three countries, it carried the wrong sign. Viewed as a whole, the estimated price elasticities given in Table 1 are similar in terms of size and statistical significance to those found in earlier studies of aggregate import behavior—see, for example, the studies by Adams and others (1969), Houthakker and Magee (1969), Taplin (1972), Samuelson (1973), and, especially, the recent survey of price elasticities in international trade by Stern and others (1975). However, in comparing the price elasticities of Table 1 with those of earlier studies, note should be taken that the studies differ in many respects including, inter alia, differences in the sample periods.[25] If, for example, our equation (1)

[23] The estimates of the parameters are asymptotically equivalent to maximum likelihood estimates.

[24] In Tables 2 and 5, the *D-W* test statistic is biased, and it should therefore be interpreted with care.

[25] That is, the study by Adams and others (1969) covers the period 1955–65, the Houthakker-Magee (1969) study the period 1951–66, the Taplin (1972) study the period 1954–70, and the study by Samuelson (1973) the period 1960–72. Further, both our study and that of Adams and others (1969) use quarterly

TABLE 1. SELECTED INDUSTRIAL COUNTRIES: EQUILIBRIUM ESTIMATES FOR TOTAL IMPORTS USING EQUATION (1), THIRD QUARTER 1955–FOURTH QUARTER 1973 [1]

Country	Constant a_0	Price Elasticity a_1	Income Elasticity a_2	ρ_1	$\bar{R}^2$	$D\text{-}W$
Belgium	−3.389	−0.624	1.746	0.20	0.922	2.02
	(5.98)	(3.42)	(14.49)			
Denmark	0.789	−0.999	0.846	0.40	0.982	1.98
	(1.18)	(5.79)	(5.88)			
France	−1.299	−1.094	1.284	0.60	0.988	2.43
	(2.14)	(4.05)	(9.96)			
Finland	−1.821	−0.319	1.415	0.20	0.969	2.06
	(9.45)	(2.83)	(36.10)			
Germany, Fed. Rep. of	−2.308	−0.699	1.516	0.40	0.995	2.15
	(4.18)	(5.05)	(12.91)			
Netherlands	−4.780	0.328	2.039	0.50	0.993	2.07
	(9.19)	(2.21)	(18.34)			
Italy	−3.749	−0.158	1.836	0.60	0.989	2.21
	(6.95)	(0.87)	(16.02)			
Japan	−1.211	0.003	1.301	0.80	0.995	1.61
	(4.01)	(0.86)	(23.25)			
Norway	−0.098	−0.812	1.019	0.10	0.980	2.01
	(0.10)	(3.23)	(5.11)			
Sweden	−1.488	−0.397	1.333	0.30	0.981	1.93
	(2.81)	(2.40)	(11.82)			
United Kingdom	−3.586	0.175	1.780	0.40	0.980	2.01
	(11.65)	(1.90)	(27.51)			
United States	−3.907	−0.448	1.837	0.70	0.987	2.51
	(7.79)	(1.97)	(17.69)			

[1] The t-values are in parentheses below the coefficients.

is estimated for the period 1955–70 rather than for the whole period 1955–73, then, as noted in Section III, the estimated price elasticities generally increase in both size and statistical significance, and they would generally be regarded as larger than those price elasticities found in most earlier studies of aggregate imports.

The estimates of the income elasticity, $\hat{a}_2$, are all positive and significantly different from zero at the 1 per cent level. With the exception of Denmark, the income elasticities tend to be large, that is, greater than unity. However, these elasticities tend to be somewhat smaller than the income elasticities obtained (say) by Houthakker and Magee (1969).[26]

data, the studies by Houthakker and Magee (1969) and by Taplin (1972) use annual data, and, finally, the Samuelson (1973) study uses semiannual data.

[26] We obtain marginally larger values for the Netherlands, Japan, the United Kingdom, and the United States. We use the Houthakker-Magee estimates as a basis for comparison of our income elasticity estimates, since their specification of the aggregate import equation is closest to our equation (1).

The fits of the equations in Table 1, as evidenced by the values of the $\bar{R}^2$, are very good. This accuracy in specification is supported by the values of the *D-W* statistic, which allows us to accept the null hypothesis of the independence of the error terms in all the estimated equations.

In the disequilibrium import equations, presented in Table 2, the "short-run" price elasticities have the expected negative signs in 8 of the 12 countries, namely, for Belgium, Denmark, France, Finland, the Federal Republic of Germany, Norway, Sweden, and the United States; in each of these 8 countries, except the United States, the estimated price elasticity is also significantly different from zero at the 5 per cent level. For most of the 12 countries, these short-run elasticities are smaller than the equilibrium price elasticities shown in Table 1. This is what one would expect.

The short-run income elasticities all carry a positive sign, and all are significantly different from zero at the 1 per cent level. As was true with

TABLE 2. SELECTED INDUSTRIAL COUNTRIES: DISEQUILIBRIUM ESTIMATES FOR TOTAL IMPORTS USING EQUATION (3), THIRD QUARTER 1955–FOURTH QUARTER 1973 [1]

Country	Constant γa_0	Price Elasticity γa_1	Income Elasticity γa_2	Lagged Imports $(1-\gamma)$	ρ_2	$\bar{R}^2$	*D-W*
Belgium	−2.130	−0.357	1.072	0.398	0.00	0.993	2.32
	(3.61)	(2.40)	(5.18)	(4.02)			
Denmark	−0.073	−0.390	0.538	0.487	0.00	0.985	2.21
	(0.13)	(2.48)	(3.30)	(4.79)			
France	−1.910	−1.230	1.678	−0.261	0.80	0.989	2.38
	(2.47)	(4.52)	(8.23)	(2.88)			
Finland	−2.036	−0.337	1.580	−0.116	0.30	0.970	2.04
	(6.99)	(2.64)	(10.01)	(1.07)			
Germany, Fed. Rep. of	−1.191	−0.271	0.721	0.547	0.00	0.996	2.30
	(2.90)	(2.35)	(4.35)	(5.46)			
Netherlands	−2.822	0.275	1.125	0.489	0.00	0.995	2.02
	(4.80)	(2.73)	(5.42)	(5.59)			
Italy	1.556	0.007	0.710	0.639	0.00	0.991	2.02
	(3.62)	(0.06)	(4.45)	(8.14)			
Japan	−0.759	0.004	0.813	0.377	0.60	0.995	2.07
	(3.88)	(0.98)	(5.65)	(3.44)			
Norway	0.051	−0.660	0.757	0.231	0.00	0.981	2.32
	(0.06)	(2.60)	(3.53)	(2.02)			
Sweden	−1.487	−0.397	1.332	0.000	0.30	0.981	1.93
	(2.72)	(2.24)	(7.27)	(0.00)			
United Kingdom	−3.772	0.180	1.849	−0.040	0.40	0.980	1.95
	(7.88)	(1.93)	(9.58)	(0.38)			
United States	−1.410	−0.171	0.664	0.640	0.00	0.988	2.29
	(3.73)	(1.79)	(4.36)	(8.24)			

[1] The *t*-values are in parentheses below the coefficients.

the estimated price elasticities, the income elasticities in Table 2 are generally smaller in absolute size than the corresponding equilibrium elasticities reported in Table 1.

The coefficient of lagged imports, $1-\gamma$, was positive and significantly different from zero at the 5 per cent level in the equations for 8 of the 12 countries. In the three countries (France, Finland, and the United Kingdom) where this coefficient assumed a negative value, probably some misspecification error is present, since such a result implies a value for γ (the adjustment parameter) greater than unity—a result that is inconsistent with one of the basic assumptions underlying the partial-adjustment model.[27]

The "long-run," or equilibrium, price and income elasticities can be calculated by dividing the short-run elasticities by γ. These long-run elasticities, along with the mean time lags in adjustment, calculated as $(1/\gamma)$, are shown in Table 3 for those 7 countries for which estimation of equation (3) yielded reasonable results.[28]

While the calculated long-run price elasticities are generally smaller than the direct equilibrium estimates shown in Table 1, the long-run income elasticities in Table 3 are generally larger. One explanation for this result could be that our use of a geometrically weighted distributed lag (on both explanatory variables) is dominated by the income variable and that a different lag structure would be appropriate for the price variable; for example, it could well be true that the proper lag pattern for the relative price variable is one characterized by an inverted *v* shape—a point to which we return in Section III. One interesting result in Table 3 is the estimated mean time lag in the adjustment of the quantity of imports to changes in the explanatory variables. In no instance was the average lag longer than three quarters. While these results by no means refute the hypothesis that the response of imports to changes in real income and relative prices is subject to recognition lags, decision lags, delivery lags, replacement lags, and production lags,[29] they do suggest that these lags may be substantially shorter than is sometimes assumed (less than two to three years). In addition, the

[27] On the other hand, in the equations for Finland and the United Kingdom (as well as for Sweden), the coefficient on log M_{t-1} was not significantly different from zero; hence, in these cases, one also cannot reject the hypothesis that γ is not significantly different from unity.

[28] We excluded a country from Table 3 if it displayed either a positive short-run price elasticity or a significant, negative adjustment coefficient in Table 2. In cases (e.g., Finland and Sweden) where the estimated coefficient on log M_{t-1} was not significantly different from zero (at the 5 per cent level), the adjustment coefficient was assumed to be unity and, thus, the short-run and long-run coefficients were assumed to be identical.

[29] See Junz and Rhomberg (1973) for a discussion of these different types of lag.

foregoing results may also explain why estimated import functions using annual or semiannual data (e.g., Khan and Ross (1975)) find little evidence of lags in adjustment for this group of countries.

The results from estimating the equilibrium import equation with the assumption that the price elasticity was itself a stochastic function of

TABLE 3. SELECTED INDUSTRIAL COUNTRIES: LONG-RUN ELASTICITIES AND AVERAGE LAGS FOR TOTAL IMPORTS, THIRD QUARTER 1955–FOURTH QUARTER 1973

Country	Price Elasticity	Income Elasticity	Average Lag[1]
Belgium	−0.593	1.781	1.7
Denmark	−0.760	1.049	2.0
Finland	−0.337	1.580	1.0
Germany, Fed. Rep. of	−0.598	2.089	2.2
Norway	−0.858	0.984	1.3
Sweden	−0.397	1.332	1.0
United States	−0.475	1.844	2.8

[1] In quarters.

TABLE 4. SELECTED INDUSTRIAL COUNTRIES: EQUILIBRIUM ESTIMATES FOR TOTAL IMPORTS USING EQUATION (5), THIRD QUARTER 1955–FOURTH QUARTER 1973

Country	Constant	Price Elasticity a_1	Income Elasticity a_2	α_1	ρ_3	$\bar{R}^2$	D-W
Belgium	−3.279 (5.71)	−0.555 (2.94)	1.723 (14.10)	−6.690 (3.54)	0.30	0.993	1.91
Denmark	0.783 (1.16)	−0.988 (5.62)	0.847 (5.86)	−0.692 (0.40)	0.40	0.982	1.99
France	−1.255 (2.02)	−1.132 (3.94)	1.275 (9.67)	0.835 (0.40)	0.60	0.988	2.42
Finland	−1.815 (9.40)	−0.280 (2.30)	1.414 (35.99)	−1.736 (0.85)	0.20	0.970	2.00
Germany, Fed. Rep. of	−2.444 (4.35)	−0.635 (4.32)	1.545 (12.94)	−1.361 (1.24)	0.40	0.995	2.14
Netherlands	−4.779 (9.13)	0.329 (2.19)	2.039 (18.20)	−0.135 (0.13)	0.50	0.993	2.08
Italy	−3.653 (6.67)	−0.161 (0.89)	1.816 (15.62)	−1.080 (1.04)	0.60	0.989	2.19
Japan	−1.292 (3.61)	0.045 (0.46)	1.318 (18.89)	−0.004 (0.43)	0.80	0.995	1.63
Norway	−0.313 (0.32)	−0.811 (3.06)	1.065 (5.10)	2.889 (1.84)	0.20	0.981	2.11
Sweden	−1.411 (2.44)	−0.365 (2.00)	1.317 (10.66)	−4.290 (1.81)	0.40	0.982	2.06
United Kingdom	−3.585 (11.61)	0.144 (1.43)	1.780 (27.41)	1.643 (0.75)	0.40	0.980	1.99
United States	−3.921 (7.78)	−0.414 (1.77)	1.840 (17.62)	−3.062 (0.62)	0.70	0.987	2.51

TABLE 5. SELECTED INDUSTRIAL COUNTRIES: DISEQUILIBRIUM ESTIMATES FOR TOTAL IMPORTS USING EQUATION (7), THIRD QUARTER 1955–FOURTH QUARTER 1973

Country	Constant	Price Elasticity $\beta_0 a_1$	Income Elasticity $\beta_0 a_2$	Lagged Imports $(1-\beta_0)$	$\beta_1 a_0$	$\beta_1 a_1$	$\beta_1 a_2$	β_1	ρ_4	$\bar{R}^2$	D-W
Belgium	−1.833	−0.471	1.115	0.294	−3.491	13.07	−13.57	13.98	0.00	0.995	2.14
	(1.90)	(2.12)	(3.23)	(1.80)	(0.06)	(0.94)	(0.66)	(1.58)			
Denmark	1.228	−0.745	0.421	0.335	−57.175	16.089	3.663	8.130	0.00	0.988	2.27
	(1.47)	(3.12)	(1.64)	(2.09)	(1.85)	(1.91)	(0.52)	(1.46)			
France	−1.690	−1.461	1.676	−0.308	−2.722	5.013	−0.576	1.207	0.80	0.989	2.34
	(1.53)	(2.98)	(5.09)	(1.73)	(0.10)	(0.47)	(0.06)	(0.24)			
Finland	−2.559	−0.296	1.796	−0.223	20.545	−7.605	−6.584	2.281	0.40	0.973	1.94
	(6.81)	(1.62)	(8.27)	(1.44)	(1.49)	(1.42)	(0.77)	(0.38)			
Germany, Fed. Rep. of	−1.235	−0.123	0.658	0.619	−3.700	−8.489	6.404	−5.573	0.00	0.996	2.32
	(1.79)	(2.60)	(2.30)	(3.51)	(0.11)	(0.71)	(0.41)	(0.53)			
Netherlands	−3.297	0.382	1.275	0.443	−5.934	−7.561	2.331	24.146	0.00	0.995	2.13
	(3.71)	(2.46)	(4.11)	(3.39)	(0.93)	(0.51)	(0.37)	(0.58)			
Italy	−0.664	−0.196	0.463	0.693	−73.460	2.766	30.797	−14.713	0.40	0.993	2.21
	(0.96)	(0.78)	(2.03)	(5.49)	(3.95)	(0.67)	(3.66)	(3.16)			
Japan	−0.848	0.042	0.839	0.370	0.206	−0.003	−0.037	0.000	0.60	0.995	2.07
	(1.37)	(0.17)	(3.79)	(3.11)	(0.06)	(0.15)	(0.59)	(0.00)			
Norway	1.632	−1.358	0.632	0.011	24.868	−2.231	11.275	−40.977	0.00	0.982	2.12
	(0.94)	(2.52)	(1.58)	(0.05)	(1.54)	(0.14)	(1.22)	(0.78)			
Sweden	−1.162	−0.522	1.309	−0.048	−26.713	−2.738	17.557	−11.503	0.50	0.983	2.01
	(1.34)	(1.81)	(5.13)	(0.30)	(1.03)	(0.25)	(1.52)	(1.21)			
United Kingdom	−2.885	−0.041	1.582	0.045	−22.690	3.878	9.960	4.925	0.50	0.982	1.98
	(3.78)	(0.26)	(5.42)	(0.30)	(0.98)	(1.44)	(0.97)	(0.89)			
United States	−3.697	−0.920	1.956	−0.164	−35.123	−0.236	12.730	−5.250	0.80	0.989	2.29
	(3.78)	(2.61)	(6.07)	(1.09)	(0.88)	(0.03)	(0.72)	(0.56)			

the change in relative prices (i.e., equation (5)) are shown in Table 4. For each of the equations, we tested for heteroscedasticity and were able to accept the null hypothesis of constant variance of the errors. In the results, only for Belgium was $\hat{\alpha}_1$ significantly different from zero at the 5 per cent level. In fact, the estimated equations are almost identical to the corresponding estimates shown in Table 1. We are thus *not* able to reject the hypothesis that the size of the price elasticity is independent of the size of the price change. Assuming constancy in this price elasticity does not appear to cause any specification bias.

The estimates of equation (7) are shown in Table 5. Recall again that if $\hat{\beta}_1$ is equal to zero, the coefficient of adjustment will be a constant, and equation (7) will reduce to equation (3). These are indeed the results that we observe in Table 5. In only one country out of 12—namely, for Italy—is $\hat{\beta}_1$ significantly different from zero at the 5 per cent level, and even then, $\hat{\beta}_1$ carries the wrong (a negative) sign. Further, as can be seen by comparing Tables 2 and 5, the overall pattern of estimated coefficients is quite similar in the two tables.[30] It therefore appears that the danger of misspecification bias owing to the assumption that importers adjust the same amount irrespective of the size of the relative price change has probably been overemphasized.

III. Further Tests of the Quantum Effect

The empirical results reported in Section II, while quite unanimous in their rejection of both versions of the quantum hypothesis, relate of course to a given set of aggregate import data, for a given time period, and for a given set of behavioral assumptions about the import demand function. In an effort to determine whether the findings of Section II could be considered as having wider applicability (i.e., whether the test results were robust), a series of additional tests of the quantum effect were conducted. More specifically, the import demand functions, including those specifications incorporating the quantum effect, were re-estimated using, in turn: (i) disaggregated import data for manufactures, (ii) data on total imports (as in Section II) but for the subperiod 1955 to 1970, and (iii) data on total imports for the period 1955 to 1973 but employing an alternative distributed-lag pattern for the real income and relative price variables.

[30] The only notable differences between Tables 2 and 5 are in the equations for Denmark, Norway, and the United States, where the estimated short-run price elasticities are much larger in Table 5 than in Table 2.

IMPORT EQUATIONS FOR MANUFACTURES

Import demand functions for total imports, like those in Section II, inevitably cover a wide variety of different goods, each of which may exhibit a somewhat different degree of responsiveness to relative price changes. It might be argued therefore that a test of the quantum effect on aggregate import data may be subject to various problems associated with the composition of price changes (across commodities)—problems that could be reduced, although not entirely avoided, by choosing a more homogeneous, disaggregated set of data. Following this line of argument, it was decided to retest the quantum hypothesis, in both of the forms outlined in Section I, on a set of quarterly import data for manufactures, that is, on commodities in the Standard International Trade Classification (SITC) groups 5–8.[31] Imports of manufactured goods were chosen for the test because earlier studies on disaggregated import behavior (e.g., see Rhomberg and Boissonneault (1964), Taplin (1972), Kreinen (1973), Khan (1975), and Stern and others (1975)) suggested that this commodity group had a relatively large price elasticity of demand—a presumably desirable characteristic for testing the quantum effect.

We were able to obtain quarterly data on the relevant variables for 8 of the 12 countries in our original sample, namely, for Belgium, France, the Federal Republic of Germany, Italy, Japan, Norway, the United Kingdom, and the United States. The sample period for most of the 8 countries was 1957 to 1973 but for a few it was shorter because of the availability of data. The specifications of the equations for manufactured imports followed closely the specifications of equations (1), (3), (5), and (7), the only differences being that the dependent variable was a volume index of manufactured imports, and that the relevant import price variable was the unit value index for manufactured imports. The domestic price (GNP deflator) and real income (real GNP) variables were the same as those used in the equations for total imports.

The estimation results for the manufactured imports equations can be summarized thus: First, for the equilibrium import equation (1), the estimated coefficient on relative prices carried the expected negative sign and was significantly different from zero (at the 5 per cent level) in the equations for five of the eight countries (namely, for Belgium, France, the Federal Republic of Germany, Japan, and Norway).

[31] Although imports of manufactured goods are often classified in the literature as covering SITC groups 5 through 9, we excluded SITC group 9 (nonclassified commodities and transactions) in order to obtain a somewhat more homogeneous commodity group.

Second, in four of the five countries where a significant price elasticity was obtained, the size of the price elasticity was larger than in the equations for total imports. This is consistent with the findings of earlier studies and is what one would expect, other things being equal. Further information on these disaggregated import equations can be obtained by consulting Table 6, in Appendix I. Third, in the disequilibrium import equation (3), four of the eight estimated price elasticities were statistically significant with the correct sign. However, the estimated adjustment coefficient performed in accord with a priori expectations (i.e., was positive and significant) in only two of the eight cases. Thus, the results of the disequilibrium model were not as a whole very encouraging for this class of imports.[32] Among other things, it could well be that the geometrically weighted lag distribution implied in equation (3) was not appropriate for manufactured imports. Finally, and most important for the purposes of this paper, estimation of both equations (5) and (7) revealed that neither of the two versions of the quantum hypothesis received any empirical support from these disaggregated import equations. More specifically, *neither* $\hat{\alpha}_1$ nor $\hat{\beta}_1$ was significantly different from zero at even the 10 per cent level in any of the eight countries. Thus, the disaggregated results are in agreement with the results for total imports as regards the quantum effect, and they do not suggest that the findings for total imports were seriously impaired by any potential aggregation bias.

TOTAL IMPORTS, 1955–70

Since behavioral relationships in economics have often proved to be unstable over time, it seems useful to determine whether the main qualitative conclusions of Section II are sensitive to the choice of sample period. To provide some information on this question, we re-estimated equations (1), (3), (5), and (7) using the quarterly data on total imports for the period 1955 to 1970.

The main results for the period 1955–70 can be characterized as follows. First, using the equilibrium import equation (1), the estimated price elasticity was negative and significantly different from zero (at the 5 per cent level) for 9 of the 12 countries. In most cases, the estimated price elasticity was larger in absolute value for the period 1955–70 than for the full period 1955–73; compare the results in Table 7 (in Appendix I) with those in Table 1. Such a result is consistent with the hypo-

[32] The results (for imported manufactures) of the estimation of equations (3), (5), and (7) are available upon application to Mr. Goldstein, International Monetary Fund, Washington, D.C. 20431, U.S.A.

thesis that relative price changes during the period 1971–73 were concentrated in those commodity groups with relative low price elasticities, thus yielding a lower aggregate price elasticity for the period 1955–73. Second, in the disequilibrium import equation (3), the estimated (short-run) price elasticities are negative for all 12 countries but are statistically significant (at the 5 per cent level) in only 6; see Table 8, in Appendix I. Further, the estimated mean time lags in adjustment are slightly longer for the period 1955–73, but the short-run and long-run price elasticities, where significant, are fairly similar as between the two periods (compare Tables 2 and 8). Third, no empirical support for the quantum hypothesis emerges from estimation of either equation (5) or equation (7). In fact, neither the coefficient $\hat{\alpha}_1$ nor $\hat{\beta}_1$ was significantly different from zero in any of the 12 countries tested. Thus, while we have not attempted an intensive investigation of the relative stability of our estimated equations over time, we find no evidence that our tests of the quantum effect in Section II were unduly influenced by the choice of sample period.

AN ALTERNATIVE LAG DISTRIBUTION

As previously mentioned, one of the limitations of the partial-adjustment model for imports, as represented by equation (3), is that it imposes the same geometrically weighted lag distribution on both the real income and relative price variables. While there appears to be general agreement in the empirical literature that the effect of real income on import demand declines rapidly over time, there is much less agreement on the proper distributed-lag pattern for relative price changes (see Leamer and Stern (1970) and Magee (1975)). For example, some writers (e.g., Heien (1968) and Samuelson (1973)) have found that the effect of relative price changes decays steadily over time with most of the response occurring during the first year, while others (e.g., Buckler and Almon (1972) and Clark (1974)) have found a bell-shaped or inverted *v* pattern for relative price changes.

In order to determine whether our earlier results for the quantum effect would be significantly altered by choosing a less restrictive lag distribution for the import function, the following two-step procedure was adopted. In the first step, we re-estimated the equilibrium import equation (1) using the polynomial (Almon) distributed-lag technique. A second-degree polynomial with a total lag period of eight quarters was selected for both the real income and relative price variables.[33] This

[33] In one case we restricted the end point in the lag distribution to be zero, while in another case we left the end point unrestricted.

procedure allows the two explanatory variables to assume different lag patterns and does not constrain the weights in each lag pattern to fall steadily over time. The results of the foregoing exercise (using the data on total imports for the period 1955–70) can be described briefly as follows: (i) the weights in the lag distribution for real income declined rather sharply over the two-year period, with most of the response falling in the first quarter for almost all 12 countries; and (ii) for the relative price variable, the lag distribution showed a more mixed pattern —in 6 countries, the weights declined steadily over time, while in the remaining 6, the weights either increased over time or took a bell shape.[34]

In the second step, we specified a new import demand equation where the real income variable was entered without a lag but where the relative price variable was permitted to affect imports with a lag. More specifically, the new import equation took the following form:[35]

$$\log M_t = c_1\lambda \log P_t^* + c_1(1 - \lambda) \log P_{t-1}^* + c_2 \log Y_t \qquad (12)$$ [36]

where M_t, Y_t, and P_t are defined as before, λ is the weight in the lag distribution for P_{t-1}^*, and where

$$P_t^* = \frac{1}{4}\sum_{i=0}^{3} P_{t-i}$$

$$P_{t-1}^* = \frac{1}{4}\sum_{i=4}^{7} P_{t-i}$$

Equation (12) expresses the quantity of imports demanded as a function of current real income and as a weighted average of relative prices in the first four quarters and the second four quarters. To test the adjustment-speed version of the quantum effect, we then express the lag distribution weight for relative prices, λ, as a function of the size of the relative price change, that is,

$$\lambda = d_0 + d_1 \,|\Delta \log P_t| \qquad (13)$$

[34] We also found that the significance of the relative price variable was appreciably lower in the polynomial import equations than in either the equilibrium or disequilibrium equations given in Tables 1 and 2, respectively. The estimated polynomial import equations can be obtained from Mr. Goldstein upon request. (See footnote 32.)

[35] The reader should be aware that the model represented in equations (12), (13), and (14) is only one (simple) possible way of testing the quantum effect while permitting different lag distributions on real income and relative prices. The difficulty, however, with alternative models that allow more sophisticated treatment of distributed lags is that they are not generally amenable to straightforward testing of the quantum effect, that is, they either involve estimation of a large number of parameters or are subject to various econometric problems.

[36] For expositional convenience, the constant and the stochastic error terms have been omitted from equations (12), (13), and (14).

Substituting equation (13) in equation (12) and rearranging terms, we obtain the following estimating equation:

$$\log M_t = d_0\, c_1 \log P_t^* + (1 - d_0)\, c_1 \log P_{t-1}^* + c_2 \log Y_t + d_1\, c_1\, |\Delta \log P_t|\, (P_t^* - P_{t-1}^*) \tag{14}$$

Note that if the response speed of imports to relative price changes is independent of the size of the relative price change, then d_1 will equal zero, and the coefficient on the last term in equation (14) will also be zero. On the other hand, if larger relative price changes do cause importers to respond more rapidly, more of the relative price response will fall in the first four quarters, d_1 will be positive and significant, and the last term in equation (14) will similarly be significantly different from zero. Therefore, estimation of equation (14) can serve as an additional test of the adjustment-speed version of the quantum hypothesis. In brief, the results indicated that λ was independent of the (absolute value of the) size of the relative price change in 11 of the 12 countries:[37] Finland was the only one where d_1 was significantly different from zero. In sum, while additional econometric work can and should be done on the treatment of lags in the import function, it does *not* appear as if our earlier reported results for the quantum effect are very sensitive to the particular lag distribution imposed on the explanatory variables.

IV. Conclusions

The purposes of this paper have been to present quarterly estimates of import demand functions for 12 industrial countries, and to empirically test the proposition attributable to Orcutt (1950) that the import price elasticity is a function of the size of the relative price change.

In the estimated equations for total imports during the period 1955–73, it was observed that import demand was responsive to relative prices for 8 of the 12 countries in our sample. The sizes of these estimated price elasticities were found to be generally similar to those price elasticities obtained in earlier studies. Real income changes were also found to exert a significant influence on aggregate import demand, and here we found that the quarterly estimates of the income elasticity tend to be somewhat smaller than do the annual or semiannual estimates.

The use of quarterly data for the relevant variables also allowed us to obtain some new estimates of the time lags in the response of aggregate imports to changes in relative prices and real income. Our results indi-

[37] The actual estimated equations for the model described earlier are available from Mr. Goldstein upon request. (See footnote 32.)

cate the presence of such lags, but the lags are substantially shorter than is sometimes assumed. For example, it was found that the average time lag in the adjustment of aggregate imports (to changes in the explanatory variables) ranges between one and three quarters. (See Table 3.) Further, there is some evidence, albeit more tentative than conclusive, that aggregate imports adjust more rapidly to real income changes than to relative price changes.

Since the size of the import price elasticity as well as the timing of adjustment are both important matters of exchange rate (or tariff) policy, our results tend to point to a favorable effect of a devaluation on the quantity of imports demanded. If our results are accurate, not only will there be a reduction in the quantity of (aggregate) imports demanded but the reduction will occur quite rapidly. Also, as long as export price elasticities of demand are not small (say, larger than 0.6 to 0.8), our results suggest that the Marshall-Lerner conditions for stability are apt to be satisfied for most of the 12 countries in our sample.[38]

We also found no evidence that either the price elasticity of demand for imports varied with the size of the relative price change, or that importers adjust any faster when faced with larger than with "normal" relative price changes. Further, these qualitative findings regarding the quantum effect emerged consistently not only in the results for total imports during the period 1955–73 but also when the tests were conducted, in turn, at a lower level of aggregation (imports of manufactures), for an alternative period (1955–70), and for an alternative distributed-lag structure. Also, Branson (1972), Laffer (1973), and Magee (1973) all reached the same negative conclusion regarding the quantum effect as did this study, but using different groups of countries, different time periods, and different (more indirect) methods of testing the hypothesis. In sum, at this point there does not seem to be any empirical support for the hypothesis that import demand will react differently (nonproportionately) according to the size of the relative price change. However, further tests of the quantum effect using more disaggregated trade data, better data on relative price changes, and alternative treatments of time lags in the import function will no doubt shed additional light on this important policy question.

[38] See Goldstein and Khan (1975) for estimates of the export price elasticity of demand for 8 of the 12 countries examined in this paper. In brief, the results reported there suggest that the Marshall-Lerner conditions would be satisfied for most of these countries.

APPENDICES

I. Further Empirical Results

TABLE 6. SELECTED INDUSTRIAL COUNTRIES: EQUILIBRIUM ESTIMATES FOR MANUFACTURED IMPORTS USING EQUATION (1) [1]

Country	Period	Constant	Price Elasticity	Income Elasticity	ρ_1	$\bar{R}^2$	D-W
Belgium	58:2 73:4	−0.061 (1.66)	−0.976 (2.09)	1.97 (6.09)	0.5	0.978	2.42
France	57:2 73:4	−11.44 (22.73)	−0.597 (2.19)	2.47 (21.97)	0.1	0.985	2.10
Germany, Fed. Rep. of	57:2 73:4	−0.068 (3.28)	−0.772 (3.63)	2.264 (11.57)	0.2	0.988	2.06
Italy	62:2 73:4	0.108 (0.43)	0.478 (1.16)	1.324 (3.69)	0.9	0.959	2.05
Japan	57:2 73:4	−0.070 (1.01)	−1.229 (3.76)	1.035 (7.34)	0.7	0.977	2.08
Norway	58:2 73:4	−0.025 (1.14)	−0.949 (1.95)	0.995 (2.50)	0.0	0.927	2.45
United Kingdom	57:2 73:4	−0.145 (7.71)	0.061 (0.31)	3.37 (38.75)	0.4	0.984	1.90
United States	57:2 73:4	−0.035 (0.92)	−0.477 (0.97)	2.896 (18.83)	0.6	0.984	2.58

[1] The t-values are in parentheses below the coefficients.

TABLE 7. SELECTED INDUSTRIAL COUNTRIES: EQUILIBRIUM ESTIMATES FOR TOTAL IMPORTS USING EQUATION (1), SECOND QUARTER 1955–FOURTH QUARTER 1970 [1]

Country	Constant a_0	Price Elasticity a_1	Income Elasticity a_2	ρ_1	$\bar{R}^2$	D-W
Belgium	−2.321 (3.38)	−0.956 (4.42)	1.519 (10.42)	0.50	0.991	1.96
Denmark	0.973 (1.54)	−1.074 (6.77)	0.806 (5.96)	0.40	0.983	2.02
France	−0.523 (0.81)	−1.310 (5.13)	1.120 (8.22)	0.70	0.986	2.46
Finland	−2.040 (6.55)	−0.247 (2.57)	1.463 (22.07)	0.30	0.967	1.94
Germany, Fed. Rep. of	−2.266 (3.40)	−0.653 (3.50)	1.508 (10.64)	0.50	0.994	2.05
Netherlands	−3.451 (6.05)	−0.382 (2.61)	1.755 (14.37)	0.40	0.994	2.08
Italy	1.039 (0.90)	−1.053 (2.88)	0.839 (3.53)	0.90	0.987	2.01
Japan	−1.778 (2.42)	−0.318 (0.98)	1.425 (9.15)	0.80	0.994	1.66
Norway	−0.567 (0.62)	−0.795 (3.34)	1.118 (5.67)	0.20	0.983	1.86
Sweden	−1.649 (2.70)	−0.427 (2.44)	1.367 (10.46)	0.40	0.984	2.15
United Kingdom	−2.919 (6.88)	0.028 (0.25)	1.637 (18.23)	0.40	0.975	2.09
United States	−4.745 (8.66)	−0.018 (1.36)	2.017 (17.89)	0.70	0.983	2.34

[1] The t-values are in parentheses below the coefficients.

TABLE 8. SELECTED INDUSTRIAL COUNTRIES: DISEQUILIBRIUM ESTIMATES FOR TOTAL IMPORTS USING EQUATION (3), SECOND QUARTER 1955–FOURTH QUARTER 1970[1]

Country	Constant γa_0	Price Elasticity γa_1	Income Elasticity γa_2	Lagged Imports $1-\gamma$	ρ_2	$\bar{R}^2$	D-W
Belgium	−1.283	−0.239	0.650	0.634	0.00	0.994	2.11
	(2.68)	(2.01)	(3.77)	(7.44)			
Denmark	−0.268	−0.370	0.579	0.488	0.00	0.986	2.26
	(0.50)	(2.41)	(3.55)	(4.58)			
France	−0.943	−1.371	1.396	−0.184	0.80	0.986	2.26
	(1.20)	(5.38)	(6.78)	(2.04)			
Finland	−2.318	−0.223	1.606	−0.081	0.40	0.967	2.00
	(5.49)	(2.29)	(8.90)	(0.71)			
Germany, Fed. Rep. of	−1.570	−0.131	0.796	0.553	0.00	0.995	2.04
	(2.88)	(3.90)	(3.90)	(5.19)			
Netherlands	−2.708	−0.175	1.196	0.395	0.00	0.994	1.98
	(5.14)	(1.70)	(6.48)	(4.45)			
Italy	−1.095	−0.145	0.606	0.643	0.20	0.990	2.07
	(1.56)	(0.71)	(2.98)	(7.32)			
Japan	−1.191	−0.204	0.929	0.354	0.60	0.994	2.00
	(1.78)	(0.72)	(4.13)	(3.03)			
Norway	−0.354	−0.581	0.792	0.283	0.00	0.983	1.94
	(0.44)	(2.52)	(3.54)	(2.14)			
Sweden	−1.395	−0.180	0.959	0.350	0.00	0.983	2.14
	(3.04)	(1.74)	(5.48)	(3.21)			
United Kingdom	−2.992	−0.173	1.935	−0.278	0.70	0.976	1.95
	(4.40)	(1.06)	(9.65)	(2.75)			
United States	−1.419	−0.021	0.606	0.701	0.00	0.984	2.30
	(3.42)	(1.36)	(3.68)	(8.73)			

[1] The *t*-values are in parentheses below the coefficients.

II. Data Definitions and Sources

In the equations for total imports, the data are quarterly for the period 1955 to 1973 and are taken from Organization for Economic Cooperation and Development (OECD), *Main Economic Indicators*, various issues. The data are seasonally adjusted and are expressed in domestic currency units at annual rates.

M = index of the volume of total imports, 1960 = 100

PM = index of the unit value of imports, 1960 = 100

PD = implicit GNP deflator, 1960 = 100

Y = real GNP, expressed as an index, 1960 = 100

$$P = \frac{PM}{PD}$$

In the equations for imports of manufactures, the data are quarterly, seasonally adjusted, and expressed in domestic currency units at annual rates. All variables are expressed as indices with 1960 = 100. Depending on the availability of data, the sample period ranges from 1957–73 for five countries to 1962–73 for Italy. The data on import values are from OECD, *Trade by Commodities* (Series B), while the data on import prices (unit value index for manufactured imports, SITC 5–8) are taken from national sources. The data for the domestic price index (the GNP deflator) and the index of real income (real GNP) were again taken from OECD, *Main Economic Indicators.*

Since, except for the United States, published quarterly series on real income or nominal income do not exist for the entire period 1955–73, it was necessary to interpolate the annual series to a quarterly basis for some years. The method of interpolation was the following. If x_{t-1}, x_t, and x_{t+1} are three successive annual observations of a flow variable $x(t)$, the quadratic function passing through the three points is such that:

$$\int_0^1 (as^2 + bs + c)\, ds = x_{t-1}$$

$$\int_1^2 (as^2 + bs + c)\, ds = x_t$$

$$\int_2^3 (as^2 + bs + c)\, ds = x_{t+1}$$

Integrating and solving for a, b, and c gives

$$a = 0.5\, x_{t-1} - 1.0\, x_t + 0.5\, x_{t+1}$$
$$b = -\,2.0\, x_{t-1} + 3.0\, x_t - 1.0\, x_{t+1}$$
$$c = 1.8333\, x_{t-1} - 1.1666\, x_t - 0.333\, x_{t+1}$$

The first two quarterly figures within any year can be interpolated by

$$\int_1^{1.25} (as^2 + bs + c)\, ds = 0.0548\, x_{t-1} + 0.2343\, x_t - 0.0390\, x_{t+1}$$

$$\int_{1.25}^{1.50} (as^2 + bs + c)\, ds = 0.0077\, x_{t-1} + 0.2657\, x_t - 0.0235\, x_{t+1}$$

and corresponding formulas give the third and fourth quarter interpolation. Multiplication by 4 expresses the interpolated series at annual rates.

We generated series on both real and nominal income and derived the implicit deflator.

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