An Aggregate Import Demand Function for Greece

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This study estimates the aggregate import demand function for Greece using annual data for the period 1951-92. There are two methodological novelties in this paper. The authors find that the variables used in the aggregate import demand function are not stationary but are cointegrated. Thus, a long-run equilibrium relationship exists among these variables during the period under study. The price elasticity is found to be close to unity in the long run. The cross-price elasticity is also found to be close to unity. Import demand is found to be highly income elastic in the long run. This implies that with economic growth, ceteris paribus, the trade deficit for Greece is likely to get worse. (JEL C22, C32, F14, O11)

Introduction

The purpose of this study is two-fold: to study the determinants of import demand in Greece using post-World War II data and to estimate the responsiveness of import demand with respect to import price, domestic price, and gross domestic product (GDP) using the recently developed methodology of cointegration.

Greece is a relatively high import country. The value of imports as a percentage of GDP has fluctuated around 30 percent in recent years. The value of exports as a percentage of GDP has been around 13 percent in recent years. Greece has faced a negative balance of trade consistently from the 1950s to the early 1990s. In 1993, exports stood at $8,777.3 million (US$), whereas imports stood at $22,759 million (US$) [Organization for Economic Cooperation and Development, 1995]. In recent years, important exports have been food, live animals, low-tech manufactured goods, clothing, and footwear. Important imports have been food, live animals, mineral fuels, lubricants, chemicals, and machinery and transport equipment.

Greece has had a long history of foreign investment and high imports. Unlike many other countries in Europe, Greece did not enjoy rapid growth in manufacturing between 1840 and 1930. Greece did not have any kind of land reform and remained a feudally-managed agrarian economy. By the end of World War II, Greece experienced large-scale capital flight. Like many other developing countries, Greece followed a protectionist policy to stem the tide of capital outflow. To preserve the value of the drachma, Greece followed restricted import policies during the 1940s and early 1950s. At the end of the 1950s, the political ideology of the Greek government lurched from parliamentary to extreme right-wing military dictatorship. With the exception of the Papandreou centrist

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administration between December 1963 and July 1965, successive Greek governments followed a largely *laissez-faire* policy on the trade front since the 1960s.

Since 1973-74, inflation has been running at high rates, making Greek products and services uncompetitive in the world market [Paleologos, 1993]. Inflation fell from 14.5 percent in 1993 to 11 percent in 1994 [Organization for Economic Cooperation and Development, 1995]. However, this rate is about 4 times the European community average. In recent years, the export of goods had picked up. However, growth was below that of world trade, suggesting losses in export market shares. Export of services, especially tourism, has picked up in the 1990s. Greece has benefitted from diversion of tourism from Mediterranean countries with political troubles. European Union transfers in the form of new infrastructure projects also continued to grow. Imports have continued to grow, and increased imports of capital goods contributed significantly to this growth. The deterioration of terms of trade also contributed to the deficit in recent years. One reason for this deterioration is the high share of low-tech products in total Greek exports. These products face increasing competition from many central and eastern European countries where wage rates are lower. Examples of such exports include leather and leather products, textiles, and clothing.

Economists have long been concerned with the estimation of trade elasticities. For policy purposes, it is important to know the determinants of import demand in Greece. All three of the elasticities measured are important. Price elasticity of import demand shows the sensitivity of import demand to changes in import prices (which can be assumed to be exogenous for a small country like Greece). A change in import price affects the real income of the trading country.

It is possible to distinguish between the two effects of a change in import price. These are the income effect and the price effect. These effects show how the real income of the country changes and how the quantity demanded for import change as a result of the change in price. In addition, there is a production effect. A rise in import price draws resources away from other industries and thus leads to an increase in the production of importables. *Ceteris paribus*, if import demand is price elastic, then a rise in the import price will lead to a decrease in the import bill and vice versa. On the other hand, if import demand is price inelastic, then an increase in the import price will increase the import bill and vice versa. Thus, if the import demand is inelastic, then a rise in the import price will make the current account deficit for Greece worse. It is clear that price elasticity of import demand is expected to be negative.

The cross-price elasticity of import demand is equally important. It shows the responsiveness of import demand to a change in the domestic price. To some extent, domestic goods and imported goods are substitutable. Thus, an increase in the domestic price is expected to increase the demand for imported goods and vice versa. The magnitude of this change in import demand has a significant impact on the current account balance.

It is equally important also to measure the responsiveness of import demand to a change in income. An increase in real income signifies growth. However, generally expect a higher import bill with such economic growth. *Ceteris paribus*, this is likely to increase the current account deficit. Again, the magnitude is very important. If the income elasticity of import demand is highly elastic (and positive), then it is likely to increase the trade deficit substantially. In this case, there is a trade-off between economic growth and the current account deficit. On the other hand, if import demand is income inelastic and small in magnitude, then an increase in income may not have a very significant effect on import demand. While income elasticity of import demand is generally expected to be positive, it may not necessarily be so. If economic growth leads to an increase in the production of goods, which would have been imported otherwise, then it is possible to have a negative income elasticity of import demand. For example, Sinha [1997] finds income elasticity of import demand to be negative for Pakistan.

Finally, the estimation of price elasticity of import demand is also important in the context of the Marshall-Lerner-Robinson condition (see Kenen [1989, pp. 294-301]). This condition states that a devaluation of a country's currency will improve the current account balance if the sum of the absolute values of the price elasticities of import and export demand of a country are greater than 1. Even though the elasticity of export demand is not estimated, a relatively high (close to 1) price elasticity of import demand will indicate that the Marshall-Lerner-Robinson condition is likely to be satisfied.

In summary, in the words of Houthakker and Magee [1969, p. 111]: "The practical and theoretical importance of price elasticities is beyond question." Surprisingly, a modern approach to the Greek import behavior is lacking and this paper attempts to fill the void. This paper uses the recently developed cointegration methodology to study the behavior of imports in Greece. Just like Houthakker and Magee, the approach here is similar in spirit to that of Neisser and Modigliani [1953] and Polak [1953].

Theoretical Issues

Import demand function estimation has a long (and controversial) history. Inspired by the balance of payment problem of the United Kingdom at the end of World War II, economists estimated the import demand function to predict whether the balance of payment was going to get worse or not (see Hinshaw [1945] and Chang [1946]). The idea was that by estimating the U.S. import demand (the largest trading partner of the United Kingdom at the time), it would be possible to see if the post-war prosperity of the U.S. could pull the United Kingdom out of their problem of balance of payment. It is remarkable to find early evidence of nonstationarity and a structural break in the time series studied by Hinshaw. Early on, researchers such as Polak [1950] have noted that estimating the demand function was facilitated by using logged variables rather than variables with deviations from mean or growth rate of variables directly. At the same time, Orcutt [1950] warned about using the ordinary least square (OLS) method to estimate the import demand function. His main objections were:

1) that supply and demand shifts happen simultaneously, therefore, it is difficult to know if what is estimated is actually tracing out a demand curve;
2) there were structural changes (in the interwar period) making estimation difficult;
3) there were errors in measuring price (index) precisely; and

4) by looking at variables contemporaneously, it is difficult to judge if the short-run or long-run effects are being estimated.

By using the Phillips-Hansen [1990] procedure, this paper effectively deals with the first problem of Orcutt [1950]. This paper also tests for a structural break to deal with the second problem. Measuring the index presents an aggregation problem, so this issue is discussed separately in this paper. By including lagged variables and estimating an error correction model, this paper addresses the fourth issue directly.

The Aggregation Problem

Demand theory relates to a particular individual. However, statistical data almost always relate to groups of individuals. An aggregation problem arises if a one-person problem cannot be extended into a group problem. There are two possible solutions:

1) interpret the demand arising from a group utility function subject to a group budget constraint; and
2) start from an individual and investigate where properties can be generalized for the individual equation carryover to the aggregate equation by summing over all individuals.

The first option is not satisfactory, as shown by Gorman [1959], because individual actions often conflict with community actions. For the second option, Wold and Jureen [1953, Ch. 7] show that if all individual incomes and all prices are modified proportionally, then all individual demands remain unchanged and, therefore, the aggregate demand, also. If aggregate income and prices are modified proportionally, then aggregate demand remains unaltered only if individual incomes vary in the same proportion. Zellner [1969] shows under what conditions that the aggregation problem does not arise. In defense of using the aggregate demand function for imports, quoting from the classic book by Houthakker and Taylor [1970, p. 200]:

> "The theory [of the dynamic preference ordering] is strictly in terms of a single individual, yet we apply it to entire countries. In doing so we ignore the aggregation problem, on which there is voluminous literature. Rather than add to this inconclusive discussion we simply state as our opinion that of all the errors likely to be made in demand analysis, the aggregation error is the least troublesome. As evidence we cite our lack of success in finding significant demographic variables, most of which would capture distributional effects. A formal discussion of aggregation would lead us too far afield, even if it were likely to be fruitful. We therefore proceed to matters of greater practical importance."

Some researchers such as Neisser [1953] strongly recommend using a disaggregated approach for estimating the import demand function. However, the problem there is finding an appropriate disaggregated price index for import. Neisser's approach is to totally eliminate the price variable and treat import demand as a function of income (or, in some cases, as industrial output). This approach is extremely unsatisfactory.

Literature Review

A number of important studies in the last three decades, which estimated the aggregate import demand function for Greece, is now reviewed. In a study of Greek imports during the period 1953-64, Sarantides [1972] finds price elasticity of aggregate imports and income elasticity of aggregate imports to be elastic. He uses the absolute price version as given in (2). Prodromidis [1975] uses the relative price version to estimate disaggregated import demand functions for many products separately as well as for total merchandise imports. Merchandise imports were found to be income elastic but not price elastic for the period 1961-69. Bahmani-Oskooee [1986], using the relative price version and an additional explanatory variable, export-weighted effective exchange rate, finds relative price elasticity and income elasticity of imports to be inelastic for Greece in both the short-run and long-run for the quarterly period 1973.I-1979.III.

There is a serious drawback to all of these studies. None of them addressed the question of stationarity of the time series. Thus, as Granger and Newbold [1974] and Phillips [1986] show, it is possible that these studies estimated spurious regressions. In this study, extensive tests of stationarity are conducted.

The Model and Econometric Methods

Import demand function can take several forms. Let M_t be imports in time t, PM_t is import price in time t, PD_t is domestic price in time t, and Y_t is real GDP in time t.

Two main forms have been used in literature:

$$M_t = f(PM_t, PD_t, Y_t) \quad , \tag{1}$$

and

$$M_t = f(PM_t / PD_t, Y_t) \quad . \tag{2}$$

Equations (1) and (2) are known as the absolute price and relative price formulations, respectively. Both forms have been used extensively in literature. The two forms have been specified in both linear and loglinear formulations. Goldstein and Khan [1985] provide an excellent summary of the earlier studies. One serious limitation of the above formulations is the assumption of instantaneous adjustments to import and domestic prices and real income on the part of the importers. Following Khan and Ross [1977], a partial adjustment model can be specified as:

$$\Delta M_t = \delta(M_t^* - M_{t-1}) \quad , \tag{3}$$

and

$$M_t^* = \alpha_1 + \alpha_2 PM_t + \alpha_3 PD_t + \alpha_4 Y_t + e_t \quad , \tag{4}$$

where Δ is a first-difference operator, that is, $\Delta M_t = M_t - M_{t-1}$, δ is the coefficient of adjustment, $0 \leq \delta \leq 1$, and M_t^* is the desired level of imports. Substituting (4) into (3) yields the dynamic linear import demand equation:

$$M_t = \delta\alpha_1 + \delta\alpha_2 PM_t + \delta\alpha_3 PD_t + \delta\alpha_4 Y_t + (1 - \delta) M_{t-1} + \delta e_t \quad . \tag{5}$$

The partial adjustment model can also be alternatively specified in loglinear form as:

$$\Delta \ln M_t = \phi[\ln M_t^* - \ln M_{t-1}], \; 0 \leq \phi \leq 1 \quad , \tag{6}$$

and

$$\ln M_t^* = \beta_1 + \beta_2 \ln PM_t + \beta_3 \ln PD_t + \beta_4 \ln Y_t \quad . \tag{7}$$

As in the linear case, substituting (7) into (6) yields:

$$\ln M_t = \phi\beta_1 + \phi\beta_2 \ln PM_t + \phi\beta_3 \ln PD_t + \phi\beta_4 \ln Y_t + (1 - \phi) \ln M_{t-1} + \phi\varepsilon_t \quad . \tag{8}$$

Equation (8) can be rewritten as:

$$\ln M_t = a_1 + a_2 \ln PM_t + a_3 \ln PD_t + a_4 \ln Y_t + a_5 \ln M_{t-1} + u_t \quad , \tag{9}$$

where $a_1 = \phi\beta_1$, $a_2 = \phi\beta_2$, $a_3 = \phi\beta_3$, $a_4 = \phi\beta_4$, $a_5 = 1 - \phi$, and $u_t = \phi\varepsilon_t$.

The coefficients of (9) will give the short-run elasticities because of its loglinear formulation. The coefficients of (7) will give the long-run elasticities. However, the coefficients of (7) can be calculated directly from the coefficients of (9) as:

$$\phi = 1 - a_5, \; \beta_1 = a_1/(1 - a_5), \; \beta_2 = a_2/(1 - a_5),$$

$$\beta_3 = a_3/(1 - a_5), \text{ and } \beta_4 = a_4/(1 - a_5) \quad .$$

Similar equations can be derived for the relative price and linear versions.

Asseery and Peel [1991] use the above model to estimate the import demand functions for 5 countries: Canada, Japan, United Kingdom, U.S., and West Germany. Doroodian et al. [1994] use a similar model to study the aggregate import demand for Saudi Arabia.

The choice between the linear and loglinear models is somewhat arbitrary. However, the Box-Cox test (see Box and Cox [1964] and Zarembka [1974]) can be used to choose between the two. This test estimates the general form of regression equation:

$$(M_t^\lambda - 1)/\lambda = b_0 + b_1[(PM_t^\lambda - 1)/\lambda] + b_2[(PD_t^\lambda - 1)/\lambda] + b_3[(Y_t^\lambda - 1)/\lambda] + \omega_t \quad . \tag{10}$$

If $\lambda = 0$, then adopt the loglinear form, and if $\lambda = 1$, then a linear function is more appropriate. Box and Cox [1964] suggest that the maximum likelihood method be used to estimate (10) to avoid the problem of serial correlation.

This study uses three types of unit root tests to check whether these variables are stationary or not. The first test is an augmented Dickey-Fuller (ADF) test, which is an extension of the Dickey-Fuller test (see Dickey and Fuller [1979, 1981]). The ADF test entails estimating the following regression equation (with an autoregressive process):

$$\Delta y_t = c_1 + \omega y_{t-1} + c_2 t + \sum_{i=1}^{p} d_i \Delta y_{t-i} + \nu_t \quad . \tag{11}$$

In (11), y is the relevant time series, Δ is a first-difference operator, t is a linear trend, and ν_t is the error term. Equation (11) can also be estimated without including a trend term (by deleting the term $c_2 t$). The null hypothesis of the existence of a unit root is $\omega = 0$.

The Phillips-Perron (PP) [1988] test is well suited for analyzing time series whose differences may follow mixed autoregressive moving average (p, q) processes of unknown order in that the test statistic incorporates a nonparametric allowance for serial correlation and heteroskedasticity in testing the regression. Consider the following equation:

$$y_t = \tilde{c}_0 + \tilde{c}_1 y_{t-1} + \tilde{c}_2(t - T/2) + \nu_t \quad , \tag{12}$$

where T is the number of observations and ν_t is the error term. The null hypothesis of a unit root is $\tilde{c}_1 = 1$. As in the ADF test, drop the trend term to test the stationarity of a variable without the trend.

The Kwiatkowski-Phillips-Schmidt-Shin (KPSS) [1992] test takes trend or level stationarity as the null hypothesis, unlike the ADF and PP tests that take the unit root as the null. The test is based on the equation:

$$y_t = c_t + c_2 t + \nu_t \quad , \tag{13}$$

where ν_t is the random error, t is the time trend as before, and c_t follows the random walk $c_t = c_{t-1} + \mu_t$ with μ_t being a random error and having a variance σ_μ^2. The null hypothesis is $\sigma_\mu^2 = 0$. As with other tests, drop the trend term in (13) if testing the stationarity of a nontrended variable.

If any variable is found to be nonstationary, this paper will test whether such a variable is stationary in its first-differenced form. Any variable that achieves stationarity after first-differencing is said to be integrated of order 1, denoted by I(1). If each variable is nonstationary but achieves stationarity after first-differencing, proceed with the cointegration tests. Two or more I(1) variables are said to be cointegrated if there exists a linear combination of them that is stationary (see Engle and Granger [1987]).

If it is found that the variables are I(1), proceed with the generalized Johansen [1991] framework of cointegration tests (see Pesaran and Smith [1998]). The general form of the vector error-correction model is given by:

$$\Delta y_t = a_{0y} + a_{1y} t - \Pi_y z_{t-1} + \sum_{i=1}^{p-1} \Gamma_{iy} \Delta z_{t-i} + \Psi_y w_t + e_t, \ t = 1, 2, ..., n \ , \qquad (14)$$

where $z_t = (y_t', x_t')'$, y_t is an $m_y x 1$ vector of endogenous variables I(1) variables and x_t is an $m_x x 1$ vector of exogenous I(1) variables:

$$\Delta x_t = a_{0x} + \sum_{i=1}^{p-1} \Gamma_{ix} \Delta z_{t-i} + \Psi_x w_t + v_t \ , \qquad (15)$$

and w_t is a $qx1$ vector of exogenous-deterministic variables, I(0) variables.

In this model, the disturbance vectors of e_t and w_t satisfy the assumptions below:

Assumption 1: $u_t = (e_t w_t)' \sim iid(0, \Sigma)$, where Σ is a symmetric positive-definite matrix.

Assumption 2: u_t (the disturbances in the combined model) are distributed independently of w_t, that is, $E(u_t|w_t) = 0$, where a_{0y} and a_{1y} (the intercept and the trend coefficients, respectively) are $m_y x 1$ vectors; Π_y is the long-run multiplier matrix of order $m_y + m$, where $m = m_x + m_y$; the Γ_{1y} and $\Gamma_{2y}, \ldots, \Gamma_{p-1,y}$ coefficient matrices capture the short-run dynamic effects and are of order $m_y x m$; and Ψ_y is the $m_y x m$ matrix of coefficients on the I(0) exogenous variables.

If the variables are cointegrated (as these variables are), then the Phillips-Hansen [1990] fully modified OLS procedure will be applied. The model is given by:

$$y_t = \beta_0 + \beta_1' x_t + u_t, \ t = 1, 2, \ldots, n \ , \qquad (16)$$

where y_t is an I(1) variable, and x_t is a $kx1$ vector of I(1) regressors, which are not cointegrated among themselves. It is also assumed that x_t has the first difference stationary process, $\Delta x_t = \mu + v_t$, $t = 2, 3, \ldots, n$, where μ is a $kx1$ vector of drift parameters and v_t is a $kx1$ vector of I(0) variables, and that $\xi_t = (u_t, v_t')'$ is strictly stationary with zero mean and a finite positive definite covariance matrix, Σ. This procedure has a number of advantages. It corrects for endogeneity and serial correlation effects, and it also asymptotically eliminates the sample bias. This procedure is applicable only where there is only one cointegrating vector.

Data and the Results

Let M_t be real import (import in billions of drachmas divided by the import price index), PM_t is the import price index (1990 = 100), PD_t is the wholesale price index (1990 = 100), and Y_t is the real GDP in billions of drachmas (in 1990 prices). Annual data for 1951 to 1992 are used and all data comes from *International Financial Statistics* [International Monetary Fund, 1996]. Hakkio and Rush [1991] argue that increasing the number of observations by using monthly or quarterly data does not add any robustness to the results in tests of cointegration. What matters more is the length of the period under consideration. Hence, the authors believe this 42-year period is sufficient for robustness of the estimates or tests.

The Box-Cox test (using the maximum likelihood method) finds $\lambda = 0.14$. The test statistic is 35.6658, and the relevant critical value of χ^2 distribution with one degree of freedom at a 5 percent level of significance is 3.84. Thus, the null hypothesis of the linear model is rejected and the loglinear model is adopted. Other studies (Khan and Ross [1977] for the U.S., Canada, and Japan and Boylan et al. [1980] for Belgium, Denmark, and Ireland) find the loglinear model more appropriate. Also, the absolute price version is used because the relative price version assumes that the influence of import price and domestic price is equal in magnitude but opposite in sign. As Murray and Ginman [1976] show, this is not borne out by empirical estimates.

Results of the ADF tests are shown in Table 1. The results indicate that the null hypothesis of a unit root cannot be rejected at the 10 percent level for any variable. Thus, the ADF tests show all 4 variables are nonstationary. Results of the PP tests are in Table 2. The results also indicate that the null hypothesis of a unit root cannot be rejected for all variables at the 10 percent level.

TABLE 1
ADF Unit Root Tests

Variables	Test Statistic	Lag Order	Critical Value
$\ln M_t$	T_τ = -2.0893	1	-3.18
$\ln PM_t$	T_μ = 0.7519	1	-2.60
$\ln PD_t$	T_μ = 1.0451	1	-2.60
$\ln Y_t$	T_τ = 0.1147	0	-3.18

Notes: T_τ and T_μ are test statistics with drift and trend and with drift and no trend, respectively. The lag order was determined using the Akaike information criterion.
Source: Critical values are from Fuller [1976, p. 373, Table 8.5.2].

TABLE 2
PP Unit Root Tests (Truncation Lag = 3)

Variables	Test Statistic	Critical Value
$\ln M_t$	-1.7796*	-3.18
$\ln PM_t$	1.0022**	-2.60
$\ln PD_t$	3.5784**	-2.60
$\ln Y_t$	0.5626*	-3.18

Notes: * denotes test statistic with drift and trend, and ** denotes test statistic with drift and no trend. The truncation lag = integer $[4(T/100)^{1/4}]$, where T is the number of observations (42 in this case).
Source: The truncation lag was determined using the Schwert [1989] criterion. Critical values are from Fuller [1976, p. 373, Table 8.5.2].

Table 3 gives the results of the KPSS unit root tests. The null hypothesis of no unit root cannot be accepted. Thus, the present model is robust as it gives the same results irrespective of the tests that are used. Extensive tests were conducted to see whether all the variables in the model are I(1), that is, whether the variables in their first differences are stationary. Both the ADF and the PP tests found each variable to be stationary in its first-differenced form. However, the KPSS tests indicated that these variables were not stationary even in their first-differenced form. Results are not reported here for brevity.

TABLE 3
KPSS Unit Root Tests (Truncation Lag = 3)

Variables	Test Statistic	Critical Value
$\ln M_t$	-1.5888*	0.119
$\ln PM_t$	1.0138**	0.347
$\ln PD_t$	1.0263**	0.347
$\ln Y_t$	2.1185*	0.119

Notes: * denotes test statistic with drift and trend, and ** denotes test statistics with drift and no trend. The truncation lag = integer $[4 * (T/100)^{1/4}]$, where T is the number of observations (42 in this case).
Source: The truncation lag was determined using the Schwert [1989] criterion. Critical values are from Kwiatkowski et al. [1992, p. 166, Table 1].

Since each variable is found to be I(1) according to the ADF and PP tests, this paper proceeds with the cointegration tests. The result of the cointegration tests are in Table 4. The results indicate that the variables are cointegrated and that the number of cointegrating vectors is equal to 1. Thus, the Phillips-Hansen fully modified OLS method can be applied to estimate (9). The Parzen window and a lag of 3 is used for estimation. The results are shown in Table 5.

TABLE 4
Trace Tests for the Import Demand Function

	Test Statistic	Critical Value
Null $r = 0$	47.3300*	47.21
Alternative $r \geq 1$		
Null $r \leq 1$	16.8100	29.68
Alternative $r \geq 2$		
Null $r \leq 2$	5.0400	15.41
Alternative $r \geq 3$		
Null $r \leq 3$	0.0708	3.76
Alternative $r \geq 4$		

Notes: * denotes significance at the 5 percent level. The critical values are at the 95 percent level. A lag of 1 was determined by using the Akaike information criterion.
Source: Critical values are from Osterwald-Lenum [1992, p. 468].

TABLE 5
Phillips-Hansen Fully Modified Estimates of (9)

Variable	Coefficient	T Ratio
$\ln PM_t$	-0.4483	-5.2543*
$\ln PD_t$	0.4825	5.7799*
$\ln Y_t$	1.3903	41.2124*
$\ln M_{t-1}$	0.4716	4.4439*
Constant	-18.0750	-18.5770*

Notes: * denotes significance at the 1 percent level.

Since the equation is estimated in log-form, the coefficients give the short-run elasticities. Thus, in the short run, price elasticity and cross-price elasticity (with respect to domestic price) of import demand are inelastic. However, income elasticity is in the elastic range. The estimated long-run elasticities are given in Table 6.

TABLE 6
Long-Run Elasticities of Import Demand

Price Elasticity	Cross-Price Elasticity	Income Elasticity
-0.8484	0.9132	2.6312

The price elasticity of demand is much higher in the long run than in the short run even though it is still in the inelastic range. This means that, other things being equal, an increase in the import price is likely to worsen the current account deficit and vice versa. Cross-price elasticity of import demand is also much higher in the long run than in the short run. Even though it is still in the inelastic range, it is fairly close to 1. Thus, an increase or a decrease in the domestic price, *ceteris paribus*, is likely to keep the import bill constant. Income elasticity of demand is much higher in the long run than in the short run. This means that an increase in income is likely to worsen the current account deficit of Greece. In other words, there seems to be a trade-off between economic growth and current account deficit in Greece. Finally, the long-run price elasticity of import demand of -0.8484 suggests that the Marshall-Lerner-Robinson condition is likely to be met for Greece. Thus, Greece can effectively use the exchange rate policies to improve its trade deficit.

Summary and Conclusions

This study estimates the aggregate import demand equation for Greece using annual data for the period 1951-92. Unlike previous studies on import demand, this study uses the cointegration methodology. This paper also tests for the appropriate form of the import demand function before estimation. The variables used in the aggregate import demand function are found to be not stationary but cointegrated. Thus, a long-run equilibrium relationship exists among these variables during the period under study. The long-run estimates do differ greatly from short-term estimates.

The long-run price elasticity is found to be in the inelastic range but not too far from unity. The long-run cross-price elasticity is even closer to unity. High long-run income elasticity implies that there is likely to be a trade-off between economic growth and the balance of trade deficit. It also implies that the persistent trade deficit problem of Greece is unlikely to get better as real income goes up. The Marshall-Lerner-Robinson condition is most likely to be met for Greece. Thus, Greece should be able to use its exchange rate policies to correct the balance of trade problem.

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