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January 2010

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ABSTRACT

This paper develops a demand function for Greece's exports of manufactures according to New Trade Theory. The sample covers a rather long period of four decades with exports aggregated based on industrial rather than on trade classification. The study contributes to a better understanding of the effects of export prices, domestic and competitors', as well as of non-price competitiveness approximated with capital stock, on export performance. The empirical estimation uses the Johansen maximum likelihood approach in the long run and a dynamic error-correction equation in the short run. The estimated long-run and short-run relationships follow the economic theory and are remarkably stable. It is shown that non-price competitiveness plays a vital role in explaining export performance in the long run as well as in the short run and that failure to include it in the export equation may lead to mis-specification error. As opposed to conventional models of export demand where income effects are very high, in the present study foreign income has a moderately high effect on exports in the long run and no effect in the short run. Exports are also sensitive to domestic and competitors' prices in the long run, but cost and price competitiveness elasticities are close to one, indicating that Greek exporters have some ability to compete on the basis of prices.

Keywords: export demand; price and non-price competitiveness; new trade theory; vector autoregressive error correction model
JEL Classification: C22; F12

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1. Introduction

The exploration of exports behaviour in small open economies has received considerable attention over the last few decades (e.g. Goldstein and Khan, 1985). Nevertheless, a significant part of empirical work on the subject has a weak theoretical background which is based essentially on Walrasian economics and unspecified theoretical models for the economies considered.

Following Goldstein and Khan (1978, 1985) the empirical work for export demand of small open economies usually contradicts theory since it detects high income and low price elasticities. Even when the export demand equation is normalized for price as in Riedel (1988), the estimated high price and low (insignificant) income elasticities prove to be due to the specification adopted (partial adjustment model). Allowing for a general dynamic specification model the estimation results are similar to those when export demand is normalized for quantity (Muscateli *etal.*, 1992). Therefore, the high income and low price elasticities that are estimated for small open economies most likely reflect omitted variables in the export demand model.

The enrichment of international trade theory with the concept of “product differentiation” opens up new scope for theoretical and empirical work. It has been shown (e.g. Venables, 1985) that when exporting firms are imperfect competitors even small open economies may be able to use policies to improve their terms of trade and to increase their welfare. According to “New Trade Theory” (e.g. Krugman, 1989), product differentiation in open economies is the most important source of trade between countries with similar economies.

For economies like Greece that face large trade imbalances a key question is whether and how export growth can contribute to reducing these imbalances in the medium to long run. The answer boils down to the likely response of exports to changes in foreign income and in price and non-price competitiveness. Solid inferences can be made about the future evolution of exports if the respective elasticities are accurate and relatively stable.

This paper has the following important features: The demand for exports function is derived according to “New Trade Theory” for a differentiated domestic good traded in foreign markets by introducing a non-price competitiveness variable. An advanced method of cointegration, is adopted to estimate empirically the demand function for Greek exports of manufactures. Trade data is aggregated based not on trade classification SITC (5-8), (which is usually adopted in most of the relevant empirical research) that categorizes about 80% of Greek manufactures, but on industrial classification (ISIC) which is more accurate and complete.[1] Finally, a long sample period and quarterly data are used in the estimation.

The main findings that emerge from this study are: First, the estimated model is remarkably stable despite the large fluctuations of exports during the rather long period (1962-1999) of the sample. Also, the speed of the short-run to long-run adjustment depends on the definition of foreign income adopted. The use of industrial production produces higher speed of adjustment compared with GDP. Second, non-price competitiveness (capital stock) plays a vital role in terms of explaining export performance in the long run as well as in the short run. Consequently, failure to include non-price competitiveness among the explanatory variables, as in most of the existing work, leads to model mis-specification. Third, foreign income elasticities, while moderately high in the long run, are not significant in the short run. Fourth, in the long run exports are highly elastic with respect to export prices (domestic and competitors') and they are at best unitary elastic with respect to unit labour cost. In the short run, price elasticities are below unity and significant while the impact of unit labour cost is not significant in most of the cases. Finally, the effect of price and cost competitiveness on exports is close to one in the long run and even in the short run. Hence, Greece has some ability to compete on the basis of prices and its exports are not completely dependent upon foreign income and non-price competitiveness.

The rest of the paper is organized as follows: Section 2 provides some background information and presents the data. Section 3 deals with the determination of the theoretical model of the demand for exports, while in Section

[1] See data section 2.2 for a more detailed discussion.

4 the econometric specification is developed. The estimation results of the long-run as well as of the short-run equations and their statistical properties are presented in detail in Section 5. Finally, Section 6 discusses the conclusions of the study. Appendix A describes the construction of the data, and Appendices B-C show the figures that accompany the econometric analysis.

2. Brief background and data

2.1 Background

During the last five decades the Greek manufacturing sector has presented a rather good export performance with an average rate of increase around 5% per year (see Figure 1). In 1962 the value of exported manufactured[2] goods financed only 11.6% of manufactured imports,[1] while the corresponding ratio rose to 50.5% in 1982 before declining to 30.5% in 2006. The role of exports of manufactured goods to the economy as a whole becomes clearer if we take into account two more statistics: Firstly, the share of exports of manufactured products in total exports more than tripled during this period (1962: 22.4%, 1982: 62.7%, 2006:71.0%). Secondly, a shift in the composition of exports of manufactured goods occurred. In 1962 exports of manufactures were dominated by exports of the processed food industry, the share of which was 50.7% of total manufactured exports, in 1982 this share dropped to 20.3% and in 2006 it was 13.9%. Industries such as, chemicals, cement and metallurgical products increased their share in exports of manufacturing.

In general, Greek exports of manufactures have shifted away from low-technology sectors towards medium- and albeit to a lesser extent high-technology sectors. The share of the low-technology sectors declined from 94% in 1960-1964 to 88% in 1985-1989 and to 67% in 2000-2001. The shares of the medium-technology and high-technology sectors increased from 6% and 0.2% in 1960-1964 to 10% and 2% in 1985-1989 and finally to 21% and 12% in 2000-

[2] Excluding exports (imports) of petroleum products.

Figure 1
Real Exports and Capital Stock in Manufacturing in Greece
(1970=100)

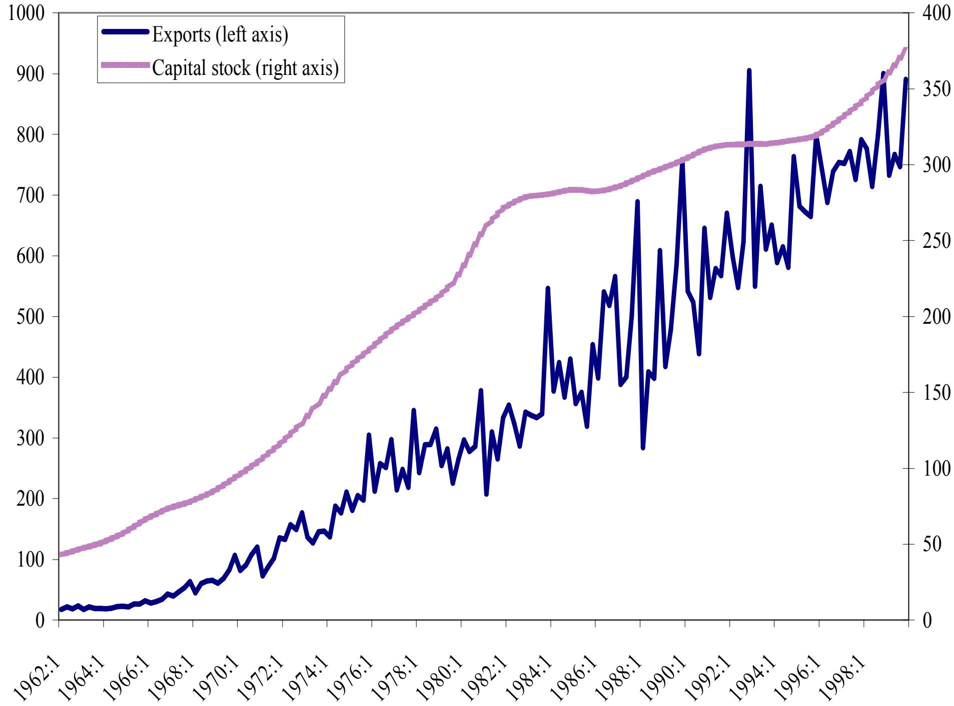
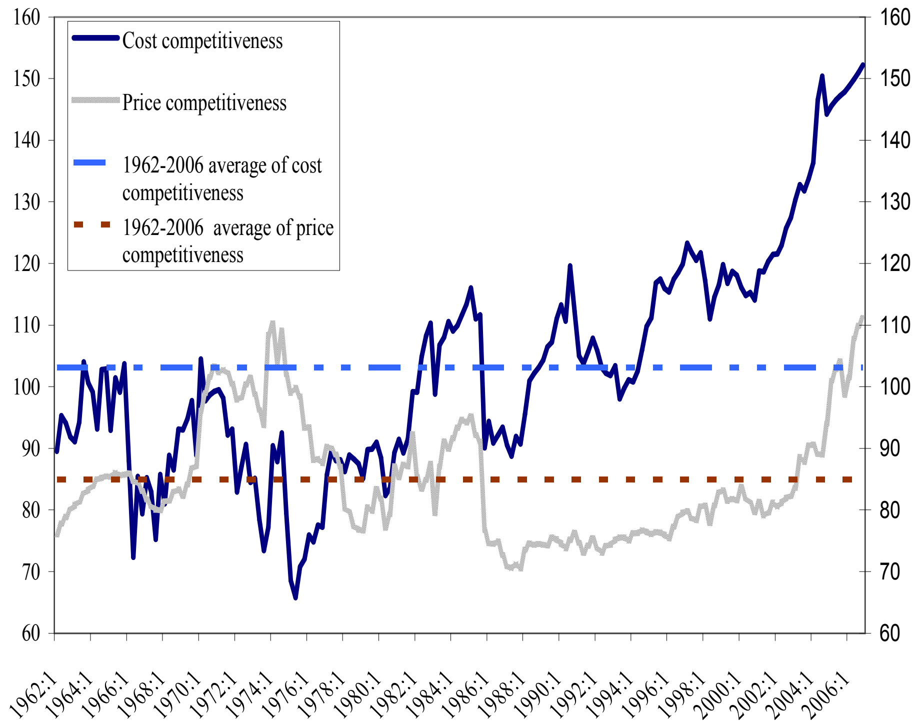


Figure 2
Export Price and Unit Labor Cost Competitiveness in Manufacturing in Greece
(1970=100, seasonally adjusted)



2001 respectively[3]. The latter shares still remain lower compared with the respective euro area average (2000-2001: 48% and 21% respectively).

Despite the good performance of exports of manufactures trade imbalances have persisted in the Greek economy and the trade deficit continued to grow in absolute terms. The trade policy followed to cope with this problem initially emphasized import substitution through high protectionism. In the late sixties and in the middle of the seventies, however, a shift in policy took place in favor of more openness. Thus, a set of measures were taken to support this policy, such as government guarantees for export finance, reduced rates or even exemption from customs duties on raw materials imported for processing and re-exporting, direct subsidies on the price of exported goods, etc. These measures aimed at improving the cost and price competitiveness of exportable goods, as well as at boosting their production.

Figure 2 presents developments in competitiveness, based on two definitions, unit labour cost in manufacturing and export prices of manufactures. During the first two decades of the sample period there were no significant differences between average cost and price competitiveness of Greek exports of manufactures and both were on average almost 11 percentage points lower than the level recorded in the base year (1970). However, from 1980 to 2006 there was a marked deterioration (deficit) in cost competitiveness. During this period, average cost competitiveness showed a significant "deficit" of 26 and 12 percentage points compared with the previous period and with the whole 1962-2006 period respectively, while average price competitiveness demonstrated a small improvement of 7 and 3 percentage points respectively. As a result, in the 1981-2006 period average cost competitiveness was almost 33 percentage points higher than the respective price competitiveness.[4] Finally, it should be noted that during the 2001-2006 period cost and price competitiveness worsened substantially, in fact the latter at a relatively faster pace. In spite of this

[3] See Bank of Greece (2005) and ECB-MPC (2005).

[4] The loss in unit labour cost competitiveness was reflected in Greece's development of the traditional labour-intensive manufacturing sectors such as footwear and clothing. During the last two decades these two sectors showed a significant cut in their production, while textile firms have partly or fully delocated their production to low-wage neighbor countries.

development, the deviation (12 percentage points) of the 2001-2006 period average price competitiveness from its long-term (1981-2006) average was lower than the respective deviation (22 percentage points) of cost competitiveness.

In contrast with these developments, other countries such as Germany, USA, Japan, Austria and Sweden have improved their cost and price competitiveness positions.[5] Accordingly, Greece's external disequilibrium can be explained to a large extent by the above mentioned "competitiveness deficit". On the other hand, from Figure 1 it is obvious that the rising investment in Greece's manufacturing sector in the 1990s should have contributed to the improvement of product quality and variety and market diversification and consequently to the increased performance of the country's exports in international markets.

2.2 The data

The data used includes the value of real Greek exports of manufactures aggregated according to the International Standard Industrial Classification (ISIC), Greek and competitors' export prices or the unit labour cost of manufactured goods, foreign income measured either by the aggregated GDP of 19 major trading partners of Greece or by OECD industrial production and the capital stock of the Greek manufacturing sector. The advantage of using the ISIC classification as opposed to SITC is twofold: First, this classification includes exports of all the manufacturing sectors, while SITC (5-8) which is often used in international trade research fails to comprise exports of industries such as processed food and beverages, tobacco, leather and furs, which during the period under examination are an important part of Greek production and exports. Second, ISIC accords with the classification method of the rest of the variables used in the estimation, thus avoiding biased coefficients in the estimation due to errors in variables. The data and sources are described in more detail in Appendix A. It should be mentioned that export prices were approximated with the wholesale price index of exports and not with the unit value of exports

[5] See OECD Economic Outlook, December 2007.

because the latter presented large measurement errors[6] during the second half of the 1990s.

3. Model specification

3.2 The framework of the model

The basic theoretical framework that is used to estimate a demand function of Greek manufacturing exports follows the "imperfect substitutes" model (Goldstein and Khan, 1985). This model assumes: First, a two country world and second, that exports are imperfect substitutes for domestic non-traded goods. Accordingly, an export demand function can be derived as the outcome of foreign households' utility maximization subject to their budget constraint. This export demand function is:

$$X^d = x\ (PX_g,\ PX_c,\ Y^f), \qquad f_1 <0,\ f_2>0,\ f_3>0, \qquad (1)$$

where X^d = the quantity of the domestic good which is exported to the foreign market, PX_g = the price of the domestic good, PX_c = the price of competing suppliers in the foreign market in a common currency, Y^f = the real foreign income and f_i = the expected partial derivatives of the export function with respect to the ith argument.

Equation (1) is a conventional export demand model although inadequate to address the effect that product differentiation has on a country's export performance. The New Trade Theory has advocated a wide stream of research on trade in differentiated products. Product differentiation can be horizontal or vertical. In the first case, pioneered by Krugman (1989), a country's relative growth increases the number of goods produced as well as the variety of products of same quality leading to the country's improved trade performance. Indeed, the observed high income elasticities for the exports of fast-growing

[6] It is well known that unit values suffer from measurement problems. The Greek series especially contain severe such problems due to the change in the classification system around mid-1990s.

economies have been attributed to the greater variety of goods produced by these economies (Krugman, 1989)[7].

However, a country's comparative advantage should be based not only on product variety but also on higher product quality (vertical product differentiation).[8] Consumers' preferences for product variants imply a preference for variety in demand subject to horizontal product differentiation. This approach of high horizontal product differentiation can describe not only final goods but intermediate inputs as well. Both horizontal and vertical differentiation can reflect either past investment choices in physical, human and knowledge capital (Thirwall, 1986) or technical improvements in the sense of moving upwards on the quality ladder, thus influencing the extent of sectoral vertical differentiation.

Despite the extensive theoretical research in this area, so far there exists only a relatively limited body of well-established empirical research about the appropriateness of the above theoretical models. Empirical knowledge about the quality and variety aspects of trade is limited, not least because of measurement difficulties. Indeed, measuring product differentiation is a challenging task since it combines a large number of products that are difficult to disentangle. There are two categories of product differentiation measures: direct and indirect. The first type uses the values or the unit values of a large set of highly disaggregated exports data[9]. Among the second are R&D expenditure, patents and investment.

We argue that investment serves as an indirect measure of product variety and quality and hence of product differentiation. Owen and Wren-Lewis (1993) found that the UK's capital stock relative to its competitors has a significant effect on its exports of manufactures. Muscatelli *etal*. (1995) have shown that non-price factors such as the capital stock can serve as a product innovation proxy and have a significant influence on NIEs exports. Madden *etal.* (1999) reached the same conclusion using investment and technology.

[7] Krugman illustrates his argument by using an increasing returns model of intra-industy trade, where there are no relative price effects.
[8] See Grossman and Helpman (1991, 1995).
[9] See Funke and Ruhwedel (2002).

In the present study, the capital stock in manufacturing is adopted as an indirect measure of product variety and quality. Therefore, equation (1) can be extended with the introduction of the non-price competitiveness proxied by the capital stock:

$$X^d = x\ (PX_g,\ PX_c,\ Y^f,\ K), \quad f_4 > 0. \tag{2}$$

The log-linearization of the explicit unrestricted form of equation (2) is:

$$X^d = x_o + x_1\ PX_g + x_2\ PX_c + x_3\ Y^f + x_4\ K + x_5\ D, \tag{3}$$

where all the variables are in logarithms and D=dummy variables to take into account changes in trade policy as well as measurement errors in our data.[10] The above equation represents the long-run cointegrating relationship among exports of manufactures and their determinants.

4. Econometric specification

It is well-established that the approach of first differencing disregards potentially important equilibrium relationships among the levels of the time series. The more recently developed method of estimating a Vector Autoregressive Error Correction Model (VECM) applying the Johansen method provides a more correct specification and is free of the simultaneous equation bias present in a single equation.

In general, the analysis of time series variables in a multivariate context involves three steps (Enders, 1995). Firstly, one has to determine the integration order of the time series, which is a prerequisite for cointegration analysis. Secondly, if the variables are integrated of the same order I(1), the next step is to estimate a long-run equilibrium relationship using cointegration analysis (Johansen, 1991 and Johansen and Juselius, 1990a). Thirdly, one has to estimate, providing that the variables are cointegrated, the dynamic behaviour of

[10] *Dint* is unit for 1971-78 and zero elsewhere taking into account changes in trade policy prior to Greece's accession to the EU in 1981, such as the gradual abolition of tariff barriers and the implementation of a subsidy scheme. D811 and D881 are unit in 1981:1 and 1988:1 respectively and zero elsewhere taking into account measurement errors.

the model which includes the residuals from the second step lagged one period as the error-correction term. The correspondence between cointegration and the Error Correction Model (ECM) is formalized in the Granger Representation Theorem (Engle and Granger, 1987).

Initially, we tested for the existence of stochastic trends in the variables. We employed the augmented Dickey-Fuller (ADF), the Phillips-Perron, and the Bierens[11] (1993) tests of the hypothesis of a unit root I(1) against the alternative of (linear trend) stationarity I(0) and the Bierens (1997) unit root test against non-linear stationarity. We also used the Bierens-Guo (1993) test of the hypothesis of stationarity against the alternative of a unit root. Finally, the above tests were used to test stationarity of the first differences of the variables.

The Johansen method is based on maximum likelihood estimation of a vector autoregressive model (VAR) of order p with Gaussian errors:

$$X_t = c_t + \sum_{i=1}^{p} A_i X_{t-i} + \Psi D_t + \varepsilon_t \,, \tag{4}$$

where $X_t = n \times 1$ is a vector that contains the five variables included in the export demand equation, $c_t = n \times 1$ vector of constants, A_i and Ψ are $n \times n$ and $n \times q$ matrices of coefficients respectively, i=1,…,p, $D_t = q \times 1$ is a vector of dummies or drift terms and $\varepsilon_t = n \times 1$ is a vector of i.i.d. errors with a positive definite variance-covariance matrix. A reparameterization of equation (4) is necessary in order to distinguish between stationarity of linear combinations of levels and of first differences. Therefore, equation (4) can be rewritten in VECM form as:

$$\Delta X_t = c_t + \sum_{i=1}^{p-1} \Gamma_i \Delta X_{t-i} + \Pi X_{t-p} + \Psi D_t + \varepsilon_t \,, \tag{5}$$

where Δ is the first difference operator, $\Gamma_i = -(I - \sum_{i=1}^{p-1} A_i)$ *and* $\Pi = -(I - \sum_{i=1}^{p} A_i)$

[11] For higher order autocorrelation.

We test the hypothesis that matrix Π can be written as Π=αβ΄, where α is a matrix representing speed of adjustment to long-run equilibrium coefficients and β is a long-run matrix of cointegrating coefficients. That is we test the null hypothesis that:

$$H_o: rank(\Pi)=r<n, \tag{6}$$

Although X_t is a vector of non-stationary I(1) variables, the relationships $\beta' X_t$ are I(0) stationary. The number of the estimated cointegrating vectors is given by the rank of Π, which is determined by the trace and the maximum eigenvalue statistic.

In the final step, the short-run dynamics were estimated, with the long-run relationship (3) entered into the short-run equation as an error correction term. In other words, it is recognized that exports do not adjust immediately to long-run equilibrium due to lags between contract and export prices, the formation of expectations and adjustment cost (i.e., transaction cost, search cost, etc.). Therefore, the single equation equivalent of model (5) can be written as follows:

$$\Delta X_t = \alpha_0 + \alpha_1 EC_{t-1} + \sum_{i=1}^{n} \beta_i \Delta X_{t-i} + \sum_{i=0}^{n} \gamma_i \Delta PX_{g,t} + \sum_{i=0}^{n} \delta_i \Delta PX_{c,t} +$$

$$+ \sum_{i=0}^{n} \varepsilon_i \Delta Y_t^f + \sum_{i=0}^{n} \zeta_i \Delta K_t + e_t, \tag{7}$$

where EC_{t-1} is the lagged error-correction term that represents the disequilibrium from the long-run relationship (3), and α_1 is the speed of adjustment coefficient. It should be noted that since we are dealing with quarterly data, equation (7) will be estimated by introducing three lags (n=3).

5. Estimation results

5.1 Preliminary estimates

In this section we empirically estimate the function of Greek exports of manufactures for the period 1962-1999, using seasonally adjusted quarterly data.

The results from all the unit root tests[12] suggest that all the variables in logs are I(1) processes. On the other hand, the first differences in logs of all the series are I(0) stationary. Four versions of equation (3) were tested for cointegration and estimated in a VECM as in (5) depending on: i) whether the export price variables (PX_g and PX_c) were represented by the export wholesale price index or by the unit labour cost index, and ii) whether the foreign income variable was the weighted average of GDP of a sample of Greece's foreign markets (Y^f_{gdp}), or the index of industrial production of the OECD countries (Y^f_{ip}).[13] The four different specifications that are shown in Table 1.B will be referred henceforth as equations 1.1-1.4. Moreover, as it will be seen below these specifications were re-estimated imposing homogeneity of degree zero to export prices in the long run or in the short run depending on where this hypothesis is accepted. These specifications are denoted as 1.1a-1.4a (see Tables 3-5).

Two important issues emerge in the process of estimation. The first concerns the formulation of the system and whether or not deterministic terms like a constant or a trend should enter the long-run and/or the short-run models. In order to answer this question apart from looking at the plots of the levels of the series, Johansen (1992b) suggests employing the "Pantula principle"[14] to have a formal test. Three different (A-C) models[15] were estimated and it was found that model A is appropriate in all equations, except for eq. 1.3 where model B should be considered. The second issue concerns the determination of the order of the VECM system. The Akaike Information Criterion (AIC) and the Hannan-Quinn (HQ) test were used in order to choose the lag length of the system for each of the four versions of equation (3). In turn, we also used an F-test by estimating[16] sequentially a VECM of lag-length of 2 up to 4. The AIC test indicated a lag-

[12] In computing the tests we have employed up to four lags to remove fourth order autocorrelation. The results from these tests are available upon request.
[13] See Appendix A for the derivation of the PXg and PXc as well as the Y^f_{gdp} and Y^f_{ip} variables. Both variables were used in Athanasoglou (1990), while Y^f_{gdp} has been used in Houthakker and Maggee (1969), Prodromidis and Anastassakou (1983) and Turnovsky (1968) among others.
[14] See Harris (1995), p. 133.
[15] Model A is the most restrictive, assuming no trends in the levels of the data and an intercept both in the long run and in the short run dynamics. Model B assumes a constant only in the short-run model. Model C is the least restrictive assuming a trend in the long-run vector.
[16] Test results are available on request.

Table 1

Cointegration analysis of export demand 1962:1-1999:4

A. Johansen's maximum likelihood ratio tests

	Export prices				Unit labour cost			
	1.1		1.2		1.3		1.4	
H_o: Number of vectors	*λtrace*	*λmax*	*Λtrace*	*λmax*	*λtrace*	*Λmax*	*λtrace*	*Λmax*
r=0 1	135.84*	98.20*	125.80*	96.09*	108.44*	79.45*	93.71*	64.61*
r≤1 2	37.64*	37.64*	29.70*	29.70*	29.00*	29.00*	29.10*	29.10*

B. Coefficients on cointegrating vector variables

	1.1	1.2	1.3	1.4
Constant	-16.389 (9.906)	-15.079 (12.714)	-14.051 (-6.933)	-12.638 (-9.228)
PX_g	-3.206 (5.888)	-2.672 (6.822)	-1.073 (-3.257)	-0.692 (-3.121)
PX_c	2.643 (5.401)	2.160 (6.118)	1.032 (2.888)	0.589 (2.379)
Y^f_{gdp}	2.619 (2.153)	-	2.296 (2.039)	-
Y^f_{ip}	-	1.792 (5.151)	-	1.045 (2.893)
K	1.311 (2.838)	1.426 (8.468)	1.281 (2.631)	1.556 (7.180)
D811[1]	-1.641 (4.009)	-	-1.796 (-5.473)	-
D881[1]	-2.564 (6.220)	-1.742 (5.866)	-1.855 (-5.558)	-0.995 (-3.695)
α_1	-0.176(7.928*)	-0.226(6.818*)	-0.296(10.117*)	-0.309(-6.457*)
α_2	-0.045(5.233*)	-0.071(5.917*)	-0.005(-0.568)	-0.059(-4.419*)

C. Joint LR test for weak exogeneity of foreign prices (costs), foreign income and capital stock

χ^2_3	2.747[0.432]	6.454[0.092]	3.703[0.295]	5.175[0.159]

D. LR test for homogeneity of export prices/unit labour cost

χ^2_1	24.10*[0.000]	38.78*[0.000]	0.409[0.522]	4.021[0.045]

Note: -The VAR is of lag order 4. The trace and maximal eigenvalue statistics are adjusted for degrees of freedom.

-t statistics are in parentheses, p values in brackets and * denotes significance at 99% level.

- PX_g is considered endogenous and the rest variables are weakly exogenous.

-The weak exogeneity statistics are evaluated under the assumption that *r*=1 in all equations.

1. D811 and D881: see footnote 10.

Table 2

Eigenvalues of the Companion Matrix.

Eq. 1.1			Eq. 1.2			Eq. 1.3			Eq. 1.4		
Real	Imag.	Mod.	Real	Imag.	Mod.	Real	Imag.	Mod.	Real	Imag.	Mod.
1.00	0.00	1.00	1.00	0.00	1.00	1.00	0.00	1.00	1.00	0.00	1.00
0.78	0.22	0.81	0.74	0.24	0.78	0.65	0.10	0.66	0.87	0.00	0.87
0.78	-0.22	0.81	0.74	-0.24	0.78	0.65	-0.10	0.66	-0.41	-0.49	0.64
-0.40	0.45	0.60	-0.41	-0.46	0.62	-0.36	-0.47	0.59	-0.41	-0.49	0.64
-0.40	-0.45	0.60	-0.41	0.46	0.62	-0.36	0.47	0.59	0.51	-0.29	0.59
-0.24	0.38	0.45	-0.26	-0.40	0.48	0.57	0.00	0.57	0.51	0.29	0.59
-0.24	-0.38	0.45	-0.26	0.40	0.48	-0.33	0.29	0.44	-0.31	0.24	0.39
0.25	0.00	0.24	0.28	0.00	0.28	-0.33	-0.29	0.44	-0.31	-0.24	0.39

length of 4 as appropriate in eq. 1.1 and a lag length of 2 in the rest three equations. The HQ test was minimized for a lag length of 2 in all equations. The F-test showed significant lags up to four quarters in all four equations. As a result, a VECM of fourth-order was chosen.

A final issue refers to the theoretical requirements of the small open economy model where foreign income and competitors' prices should be considered as exogenous, while the non-price competitiveness variable as predetermined. In this context, individual and joint weak exogeneity tests were carried out for these variables.[17] Table 1.C reports the likelihood ratio (LR) test statistics testing for the joint weak exogeneity hypothesis. It can be seen that this hypothesis was not rejected for the above mentioned variables in all four equations.[18] Thus, weak exogeneity of these three variables was imposed on the estimation of the VECM model.

Table 1.A presents the estimated trace (λ_{trace}) and maximum eigenvalue (λ_{max}) test statistics that determine the reduced rank of matrix Π and therefore the

[17] See Johansen (1992a) and Engle and Hendry (1993).

[18] Similarly, the LR tests for individual weak exogeneity did not reject this hypothesis for the three variables. The results are available upon request.

number of cointegrated vectors.[19] The statistics indicate the presence of two cointegrating vectors in all four specifications based on 1% level of significance critical values[20] from Pesaran *etal*. (2000). [21] Thus, the full rank hypothesis is accepted. Observing that the first vector has coefficients with the anticipated sign and statistically significant and knowing that we have used dummy and weakly exogenous variables, which according to Juselius (1995) "is likely to change the asymptotic distributions to some (unknown) extend" we conclude that we need additional evidence to determine the cointegration rank. In order to choose the number of r we consider the companion matrix of the system and its characteristic roots. This cointegration test determines the number of the common stochastic trends, in other words the cointegration rank of the system, by the number of the roots of the companion matrix that are close to unity. Table 2's results reveal the existence of one common stochastic trend (one significant long-run vector[22]) in all four equations, since in all cases there exists only one unity-valued characteristic root and the second is significantly below unity ("adequately small").

5.2 Empirical results

5.2.1 Long-term estimates

Four main points emerge from the examination of Table 1.B. First, the estimated coefficients take the theoretically expected sign and are highly significant.[23] Second, exports are highly elastic with respect to domestic export

[19] The trace and maximum eigenvalue tests presented in this study are adjusted for small samples, to avoid overrejection of the null of no cointegration, which occurs due to small sample bias (see Reimers, 1992).

[20] These critical values should be used with caution since the estimation process of the econometric model involves dummy variables.

[21] It should also be borne in mind that both statistics test the same null hypothesis of cointegration rank *r* (the number of cointegration vectors) but the trace statistic's alternative hypothesis tests that cointegration rank is larger than *r* while the maximal eigenvalue statistic's alternative hypothesis tests that the cointegration rank is *r*+1.

[22] In addition, the first eigenvalue is substantially higher than the remaining four, which is evidence in favor of one cointegrating vector.

[23] In eq. 1.3 a constant was included in the cointegration space, since it was found to be highly significant, despite the outcome of the "Pantula principle" that suggested as appropriate for this equation Model B, according to which the constant should be present only in the short run dynamics and not in the long-run vector.

Table 3

Imposing homogeneity of prices in the long run and in the short run export demand

Equation / Variables	Long-run		Short-run	
	1.3a	1.4a	1.1a[1]	1.2a[1]
Constant	-13.268 (-8.271)	-10.443 (-11.139)	-	-
Lagged exports	-	-	-0.395 (-6.007)	-0.352 (-5.004)
Price/cost Competitiveness	-1.169 (-3.697)	-0.934 (-4.223)	-1.139 (-5.399, -2.483)	-1.089 (-5.367, -2.099)
Foreign income	2.162 (1.946)	0.808 (2.124)	1.055 (1.476)	0.587 (1.651)
Capital stock	1.265 (2.517)	1.466 (6.254)	2.766 (6.096)	2.513 (5.636)
EC_{t-1}	-	-	-0.069 (-3.442)	-0.115 (-3.188)
Dint[2]	-	-	-0.044 (-4.931)	-0.032 (-3.209)
D811[2]	-1.857 (-5.486)	-	-0.455 (-5.916)	-0.415 (-5.929)
D881[2]	-1.942 (-5.654)	-1.163 (-3.909)	-0.463 (-6.035)	-0.445 (-6.095)
ρ_2	-	-	-0.250 (-2.638)	-0.260 (-2.656)
Log likelihood	-	-	148	149
SER	-	-	0.09	0.09
Jarque-Bera $\chi^2(2)$	-	-	0.517[0.772]	0.462[0.793]
F(ARCH(4))	-	-	0.142[0.966]	0.202[0.936]
F(RESET)	-	-	0.564[0.570]	1.094[0.338]
F(HET)	-	-	0.814[0.756]	0.518[0.987]
LM(4)	-	-	0.718[0.581]	1.160[0.331]

Note: t statistics are reported in parentheses and p-values in square brackets.

1. In the short-run equations 1.1a and 1.2a, the coefficient of competitiveness is found by summing the coefficients of the level and the first lag of relative prices. Relative prices are expressed as the ratio of the first lag of Greek export prices over the level of competitors' prices.
2. *Dint,* D881, D811: see footnote 10.

prices (-3.2 in eq. 1.1 and -2.7 in eq. 1.2) and competitors' prices (2.6 in eq. 1.1 and 2.2 in eq. 1.2) and the respective elasticities are more than twice to even four times the unit labour cost ones (-1.1 and 1.0 in eq. 1.3 and -0.7 and 0.6 in eq. 1.4 respectively). However, the elasticity of domestic export prices is higher than that of Greece's competitors implying that the impact of price competitiveness on exports is higher than unity, while cost competitiveness is close to unity. Thus, in the long run Greek exports of manufactures are quite sensitive to price changes.

The hypothesis of long-run export price homogeneity of zero degree is relevant for export demand. Price homogeneity can be justified on microeconomic grounds, but the extent to which it applies in an aggregate trade model such as the one described in the present study should be established empirically.[24] This hypothesis tests the restriction of whether the coefficients of domestic and competitors' prices are equal and with opposite sign. The restriction imposed is overidentified, since we have already shown that there is one cointegration vector.[25] The results are reported in Table 1.D. The LR test statistics reject the null hypothesis of zero degree homogeneity in eqs. 1.1 and 1.2 and fail to reject it in eqs. 1.3 and 1.4, where export prices are defined by unit labour cost. The overidentified cointegration vectors (eqs. 1.3a and 1.4.a) are presented in Table 3, where it is shown that the elasticity of exports with respect to cost competitiveness is close to unity and highly significant.

Furthermore, in Appendix B Figures B5-B8 plot the recursive LR statistic and the respective critical value at 1% level of significance for the four equations. These figures verify that the restriction is rejected for eqs. 1.1 and 1.2 for a large part of the sample, while it is not rejected for eqs. 1.3 and 1.4. Thus, the above

[24] It is common practice to assume absence of money illusion and to introduce the ratio of the two prices in the export function. Ahluwala and Hernadez-Cata (1975) discuss extensively the invalidity of the assumption of zero degree homogeneity of prices in an export demand equation. Athanasoglou (1990), Arize (1987) and Wilson and Takacs (1979) are among the few who test export price homogeneity.

[25] Identification of the cointegration space requires imposing r(r-1) (where r represents the number of significant cointegrating vectors) linearly independent restrictions, after normalizing the cointegrating vectors. Restrictions over this number are overidentified.

results indicate that the a-priori imposition of the price homogeneity restriction, which is observed in other studies, can produce unstable estimates.[26]

Third, when foreign income is defined by GDP the corresponding elasticities (2.6 in eq. 1.1 and 2.3 in eq. 1.3) are quite higher than those when it is defined by industrial production (1.8 in eq. 1.2 and 1.0 in eq. 1.4). These elasticities are slightly lower when relative unit labour cost is considered (see Tables 3 and 4). From the magnitude of the estimated elasticities it can be seen that foreign income has a moderately high effect on Greek exports. In addition, these elasticities are similar or even smaller than those obtained in studies of export demand for other countries. For instance, an average foreign income elasticity of around 2.50 for 8 industrial countries was found by Goldstein and Khan (1978) for total exports, while Muscatelli *etal.* (1992) estimated higher long-run income elasticities, around 4, for the newly industrialized economies.

Fourth, in all equations including the homogeneous case exports are elastic with respect to the capital stock variable with robust coefficients ranging from 1.3 to 1.6. As has been mentioned before, the omission of the non-price competitiveness variable from the traditional model of export demand results in mis-specification bias in the coefficients of price and foreign income variables. To verify this argument eqs. 1.1-1.4 were re-estimated excluding capital stock. Table 4 compares long-term export price and foreign income elasticities when capital stock is included and when it is not for each of the four equations. Price (domestic and competitors') and mainly foreign income elasticities are markedly lower when non-price competitiveness (capital stock) is included among the explanatory variables.[27] It is notable that this behaviour can also be observed even under the export price homogeneity restriction in the long-run eqs. 1.3a and 1.4a.

Following these results it seems that non-price competitiveness that is represented in this model by capital stock has a significant effect on manufacturing exports elasticities in Greece. If its impact is not taken into

[26] See Catao and Falcetti (2002), Chinn (2004) and O'Donnell (2005).
[27] It is worth mentioning that in Muscatelli *etal.*(1995) it is found that these elasticities are reduced.

Table 4
Long-run and short-run price and income elasticities of export demand with and without capital stock

Elasticities	Equations	1.1		1.2		1.3		1.3a		1.4		1.4a	
		with capital stock	without capital stock	with capital stock	without capital stock	with capital stock	without capital stock	with capital stock	without capital stock	with capital stock	without capital stock	with capital stock	without capital stock
Long-run	Domestic prices	-3.206	-4.133	-2.672	-5.838	-1.073	-1.490	-	-	-0.692	-1.459	-	-
	Competitors' prices	2.643	3.334	2.160	4.759	1.032	1.468	-	-	0.589	1.428	-	-
	Relative prices	-	-	-	-	-	-	-1.169	-1.538	-	-	-0.934	-1.512
	Foreign income	2.619	6.307	1.792	4.643	2.296	5.170	2.162	5.080	1.045	3.151	0.808	3.034
Short-run	Domestic prices	-0.358	ns	-0.358	ns	ns	ns	ns	ns	-0.585	-0.592	-0.541	-0.502
	Competitors' prices	0.558	0.667	0.613	0.753	ns	ns	ns	ns	ns	ns	ns	ns
	Relative prices	-1.139[1]	-0.571	-1.089[2]	-0.483	-	-	-	-	-	-	-	-
	Foreign income	ns	1.365	ns	ns	ns	3.139	ns	3.257	ns	1.063	ns	0.950
	Foreign income in relative price equation	ns[1]	4.074	0.587[2]	1.299	-	-	-	-	-	-	-	-

Note: ns denotes non-significant, exceeding the 10% level of significance.
1. Results from eq. 1.1a, Table 2.
2. Results from eq. 1.2a, Table 2.

consideration, the resulting estimates of income and to a lesser extent of price elasticities are biased and show an inflated effect which incorporates the non-price factors like product variety and quality. Thus, the non-inclusion of the non-price competitiveness variable in the export demand equation renders the model mis-specified.

Finally, Table 1.B reports the coefficients α_1 and α_2 of the long-run model that represent the speed of adjustment to long-run equilibrium of the two equations, namely exports and domestic export prices. We observe a relatively fast adjustment in the export demand equation, while the adjustment in the domestic export price equation is smaller but not negligible. Thus, there is some gain in efficiency of the estimators if domestic export prices are considered as endogenous which was done in the present study.

5.2.2 Short-term estimates

In the next step, the short-run export equation (7) was estimated with maximum likelihood (ML) using up to fourth order autoregressive and moving average errors.[28] The representative specification of eqs. 1.1-1.4, without any restriction, and of 1.3a-1.4a where the error correction term contains the homogeneity restriction (since the hypothesis was accepted in the long-run), was found by applying “the general to specific” methodology of Davidson *etal.* (1978), and Hendry (1987), among others. Results are summarized in Table 5 along with a battery of tests on the statistical adequacy of the models.[29] Given that the homogeneity of export prices hypothesis was accepted in the short-run equations 1.1 and 1.2 they were re-estimated in restricted form and the results are presented in Table 3 as eqs. 1.1a and 1.2a.

[28] In the OLS estimation (the results are not presented) the LM(4) test rejected the presence of serially correlated errors. However, the ML method with autoregressive errors shows that in all equations the fourth order and in eq. 1.1 the second order serial correlation are significant (see Table 5).
[29] In all equations the constant term appears to be insignificant and therefore is not presented in Table 5.

Table 5

The short-run error correction representation of export demand
1962:1-1999:4

Equations / RHS variables	Export prices		Unit labour cost			
	1.1	**1.2**	**1.3**	**1.3a**	**1.4**	**1.4a**
Constant	-	-	-	-	-	-
ΔX^{d}_{t-1}	-0.522 (-10.254)	-0.340 (-5.901)	-0.316 (-4.625)	-0.297 (-4.511)	-0.271 (-4.159)	-0.251 (-3.834)
$\Delta PX_{g,t-1}$	-0.358 (-2.696)	-0.358 (-2.139)	-0.186 (-0.745)	-0.259 (-1.027)	-0.585 (-2.380)	-0.541 (-2.341)
ΔPX_{c}	0.557 (3.503)	0.613 (3.581)	-	-	0.004 (0.029)	0.039 (0.289)
$\Delta PX_{c,t-1}$	-	-	-0.025 (-0.164)	0.022 (0.149)	-	-
ΔK_{t}	-	1.412 (2.924)	-	1.339 (2.313)	-	-
ΔK_{t-2}	1.943 (4.904)	-	1.716 (2.753)	-	1.247 (2.567)	1.348 (2.866)
$\Delta Y^{f}_{gdp,t-1}$	0.867 (1.368)	-	-	-	-	-
$\Delta Y^{f}_{gdp,t-3}$	-	-	0.944 (0.824)	1.281 (1.113)	-	-
ΔY^{f}_{ip}	-	-	-	-	0.624 (1.439)	0.484 (1.127)
$\Delta Y^{f}_{ip,t-2}$	-	0.239 (0.540)	-	-	-	-
EC_{t-1}	-0.050 (-1.656)	-0.162 (-3.665)	-0.035 (-1.782)	-0.042 (-2.317)	-0.173 (-3.985)	-0.171 (-4.275)
D811[1]	-0.526 (-6.754)	-0.469 (-5.584)	-0.476 (-4.391)	-0.635 (-6.854)	-0.464 (-4.906)	-0.458 (-4.898)
D881[1]	-0.468 (-6.265)	-0.637 (-8.333)	-0.623 (-6.428)	-0.461 (-4.375)	-0.631 (-6.802)	-0.639 (-6.951)
ρ_2	-0.293 (-3.159)	-	-	-	-	-
ρ_4	-	-0.267 (-2.041)	-0.359 (-4.402)	-	-0.348 (-4.253)	-0.353 (-4.319)
Log –likelihood	146	144	130	135	137	138
SER	0.09	0.09	0.10	0.10	0.09	0.09
Jarque-Bera $\chi^2(2)$	0.697[0.705]	1.494[0.473]	0.029[0.985]	0.102[0.950]	0.156[0.925]	0.230[0.891]
F(ARCH(4))	0.509[0.728]	0.705[0.589]	0.322[0.862]	0.334[0.854]	0.243[0.913]	0.231[0.921]
F(RESET)	0.544[0.703]	0.004[0.999]	0.904[0.464]	1.165[0.329]	1.068[0.375]	0.733[0.571]
F(HET)	0.913[0.597]	0.676[0.887]	0.849[0.706]	0.594[0.947]	0.831[0.711]	0.753[0.809]
LM(4)	1.608[0.175]	1.060[0.378]	2.700[0.033]	1.358[0.252]	0.643[0.633]	0.499[0.736]
Wald test for homogeneity of export prices/unit labour cost						
$\chi^2(1)$	1.557[0.214]	1.870[0.174]	-	-	-	-

Note: Δ denotes first differences, t statistics are in parentheses and p values in square brackets. The ML method of estimation, which was used, applied up to fourth order autoregression, where the second order (ρ_2) coefficient in eq. 1.1 and the fourth order (ρ_4) coefficient in eqs. 1.2, 1.3, 1.4 and 1.4a were found to be significant. Moving average errors were also added where they were found significant. SER is the standard error of the regression; Jarque-Bera is the chi-square normality test of residuals, F(ARCH(4)) is the F test for 4th order autoregressive conditional heteroscedasticity, F(RESET) is the F test for first order Ramsey's test for specification error, F(HET) is White's test for heteroscedasticity and LM(4) is the LaGrange Multiplier F test for 4th order serial correlation.

The estimated coefficients that represent short-run impact elasticities, have the signs predicted by theory. The coefficient of the error-correction term is negative and statistically significant supporting the cointegration hypothesis. This coefficient is higher than that identified for other countries (see Chinn, 2003 and O'Donnell, 2005). Adjustment to disequilibrium is relatively slower (4%-7%) when foreign income is represented by GDP and faster (12%-17%) when foreign income is defined by industrial production. In other words, the adjustment is completed at most within 6 years in the former case and in two years in the latter. Since industrial production reflects short-run adjustments in the production process it produces higher speed of adjustment compared with GDP. On the other hand, GDP by construction (it was generated from annual data) is more of a long-term variable and contributes to a slower adjustment.[30] Foreign income has an insignificant effect on exports in the short run.

Domestic and competitors' export prices have a significant but small effect on exports in all equations, but domestic unit labor cost is significant in eqs. 1.4 and 1.4a only. Price homogeneity was not rejected in eqs. 1.1 and 1.2 as the Wald test shows (see Table 5). Given this result, these equations were re-estimated in restricted form and the estimates are presented in Table 3 as eqs. 1.1a and 1.2a. It is shown that in the short run price competitiveness is significant with the right sign and close to unity.

The capital stock appears to have the strongest effect. The short-run elasticities are between 1.2 and 2.8 in all eight estimated equations and highly significant. A higher short-run than long-run capital stock coefficient emerges due to the nature of this variable, being stock in the long run and flow in the short run. The short-run eqs. 1.1-1.4a were re-estimated excluding capital stock (as it was done in the long-run equations) to test for mis-specification bias (see Table 4). In the unrestricted (non-homogeneous) equations foreign income becomes insignificant when non-price competitiveness (capital stock) is included among the explanatory variables. By contrast, domestic export prices become significant but with low elasticities. On the other hand, in the restricted equations 1.1a and

[30] See also Chinn (2004).

1.2a relative export price elasticities from inelastic increase slightly above unity. The above result implies that when the short-run export demand function is mis-specified, i.e. by not including the non-price competitiveness variable, export price elasticities tend to be low as is found in other studies (Prodromidis, 1975, for Greece, Goldstein and Khan, 1978, for industrial countries and Muscatelli *etal*., 1995, for Asian economies).

5.2.3. Statistical Properties of the Long-Run and the Short-Run Models

In light of the results of the estimation of the long-run and the short-run export demand equations, this section presents their statistical properties consisting of diagnostic tests of performance, goodness of fit tests and stability tests especially desired for the purpose of policy inferences.

The diagnostics of the short-run equations shown in Table 5 indicate a satisfactory performance. The standard errors of the regressions are not high, considering that the variables are expressed in first differences. The residuals are normally distributed, with no serial correlation and autoregressive conditional heteroscedasticity. In addition, there is no evidence that the estimated equations are mis-specified.

Another critical issue of model performance is the closeness of the fitted values of the models to the actual series. Table 6 shows three statistics for model evaluation, namely the root mean square error, the percent mean square error and Theil's inequality coefficient. As it is known the smaller these statistics are the better is the model. In the long-run eq. 1.2 shows the best performance and in the short-run eqs. 1.1 and 1.2 present the best fit. In addition, plots of the actual series and fitted (short-run and long-run) values, that are available upon request, show the ability of the model to reproduce the turning points in the data.[31]

[31] Eq. 1.2 records the smallest deviations of the fitted values from actual data and this can be related to the high value of the error-correction coefficient in this equation. Thus, the goodness of fit of these equations does not depend on the definition of the competitiveness measure but rather on foreign income defined by industrial production.

The stability of the estimated long-run vector has to be established to ensure robustness of the model. Stable cointegrating vectors are associated with a smooth time trend path of the eigenvalues. We estimated recursive eigenvalues associated with the first eigenvector that correspond to eqs. 1.1-1.4, and are presented in Figures B1 to B4 of Appendix B. All four equations do not reveal any evidence of parameter instability with the exception of eq. 1.3 which has a structural break in the first half of 1970's. A further investigation with the

Table 6

Model evaluation statistics for the long-run and the short-run model

	Export prices				Unit labour cost			
Equations	**1.1**	**1.1a**	**1.2**	**1.2a**	**1.3**	**1.3a**	**1.4**	**1.4a**
Long-run model								
RMSE[1]	0.298		0.203		0.555	0.573	0.260	0.252
RMSPE[2]	0.030		0.020		0.053	0.055	0.026	0.025
Theil's inequality coefficient[3]	0.022		0.015		0.040	0.041	0.019	0.018
Short-run model								
RMSE	0.178	0.404	0.206	0.304	0.274	0.205	0.263	0.227
RMSPE	0.018	0.040	0.020	0.030	0.027	0.021	0.026	0.023
Theil's inequality coefficient	0.009	0.019	0.010	0.015	0.013	0.010	0.013	0.011

Note:

1. Root mean square error (RMSE) = $\sqrt{1/T\sum_{t=1}^{T}(Y_t^s - Y_t^a)^2}$,

2. Root mean square percent error (RMSPE) = $\sqrt{1/T\sum_{t=1}^{T}((Y_t^s - Y_t^a)/Y_t^a)^2}$,

3. Theil's inequality coefficient = $\dfrac{RMSE}{\sqrt{1/T\sum_{t=1}^{T}(Y_t^s)^2} + \sqrt{1/T\sum_{t=1}^{T}(Y_t^a)^2}}$,

where Y_t^s = fitted value of exports, Y_t^a = actual value of exports, T=number of periods in the simulation. The actual and fitted values are in logarithms.

recursive beta coefficients shows some instability only in eq. 1.3 in the beginning of 1980's.[32] In conclusion, in the long run the estimated cointegration vectors that correspond to eqs. 1.1, 1.2 and 1.4 appear to be temporally stable.

Regarding stability of the short-run equations the following tests have been applied: 1) the cumulative sum of the squared residuals (CUSUMSQ), 2) Quandt's likelihood ratio test[33], and 3) the Chow test. In Appendix C, Figures C1-C8 present Quandt's test, while results from the CUSUMSQ and Chow tests are available upon request. The CUSUMSQ test does show some instability in the period 1977-1981[34] for eqs. 1.1 and 1.1a, while the remaining eqs. 1.2-1.4a do not exhibit any significant instability at 5% level of significance. For the lowest points in Quandt's test and the significant breaks in CUSUMSQ tests, the Chow breakpoint test was applied and it accepted the stability hypothesis at 5% level of significance in all equations.

The question then is whether there is any ground on which to choose the best from the alternative specifications. Based on the evaluation of model performance (sign, coefficient size, significance, interpretability, diagnostics, plots of fitted and actual values, evaluation statistics and the stability tests) eq. 1.2 where export prices are not treated as homogeneous and industrial production is the foreign income variable is the best specification both in the long run and in the short run. Thus, eq. 1.2 is the best approximation to the data generating process among all the estimated versions of equations (3) and (7), while its short-run homogeneous version, eq. 1.4a, offers a second best approximation.

6. Conclusions

In this paper the demand for Greek exports of manufactures was estimated using the framework of "New Trade Theory" which suggested augmenting the traditional model in which exports is a function of prices (domestic and competitors') and foreign income, with a non-price

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[32] Results are available upon request.

[33] See Quandt (1960).

[34] The detected instability in the mid-1970s may reflect the enrichment of export promoting trade policies pursued during that time with a new scheme of direct subsidies on exports of manufactures.

competitiveness variable proxied by the capital stock. The analysis assumes imperfect competition in international markets, where trade consists mainly of exchange of differentiated products. Thus, even in a small country, firms produce and export their unique varieties and can influence export prices. The empirical estimation is performed using multivariate cointegration analysis for a sample covering a rather long period using data according to the ISIC. The cointegration econometric methodology allows for the determination of a long-run equilibrium relationship and in a second stage, by imposing the long-run result, it estimates a short-run adjustment to equilibrium.

The econometric results are in accord with the predictions of the theoretical model. Both long-run and the short-run export demand equations are remarkably stable in three out of four cases and they exhibit satisfactory in-sample predictive performance. In the preferred short-run equation adjustment to long-run equilibrium is rather fast and it is completed within a year and a half. An important result pervasive in all specifications is that non-price competitiveness, often omitted from conventional models of export demand, is crucial both in the long run and in the short run and it has a direct strong positive effect on export performance as well as an indirect one via reducing the effect of export prices and price competitiveness. Another important result of our estimation is that foreign income determines export demand in the long run but not in the short run. The long-run elasticity is moderately high and close to that found in related research for other industrial countries. Long-run and short-run foreign income elasticities are also higher when GDP is the foreign income variable.

Further, export prices (domestic and competitors') have a strong effect on exports in the long run while their short-run impact is moderate. On the other hand, unit labour cost appears to be significant only in the long run with elasticities higher when GDP is the foreign income variable. Also, long-run export price elasticities are higher compared with those of unit labor cost. Thus, we could argue that to broader price or income definitions correspond higher estimates of long-run elasticities. Regarding export prices, one possible explanation of the above behaviour is that exporters base their decision making to export on prices rather than on cost. Another one, is that price determination by firms (mainly the oligopolistic ones) takes into account “standard cost” (a long-run variable) rather than actual and a variable mark-up. Homogeneity of prices is accepted in the long run when they are measured by unit labour cost and in the short run by export prices and the elasticity of the respective measures of competitiveness was found to be close to unity. This is evidence that Greek exporters face rather moderate competitive pressure in the foreign markets.

The above empirical findings have clear policy implications. Policy measures should include innovation promoting activities, improvements in product variety and quality, the creation of a more efficient investment environment and the increase of investment and R&D expenditure. In general, policy measures should halt the erosion of Greece’s international price and mainly cost competitiveness.

APPENDIX A

Data definition and sources (for more details see Athanasoglou, 1990) are as follows:

X^d = the value (in constant 1990 prices, million drachmas) of exports of manufactures according to the International Standard Industrial Classification (ISIC), excluding exports of petroleum products, dried fruits, manufactured tobacco and ginned cotton. Trade data according to SITC reclassified according to ISIC. Source: "Foreign Trade Statistics", National Statistical Service of Greece (NSSG).

PX_g = the wholesale price (or unit labour cost index) of Greek exports of manufactured goods, 1990=100. Source: NSSG.

PX_c = wholesale or unit value (or unit labour cost) index of exports of manufactured goods of Greece's competitors, 1990=100. This index is derived according to the following formula:

$$PX_c = \sum_{i=1}^{n} (PX_i)\, W_i$$

where i = 1, ….. n Greece's competitors, PX_i = export prices (or unit labour cost) of Greece's n competitors, W_i = the n weights

The weights Wi's were borrowed from Durand's (1986) study for the period 1960-75. For 1975-99 OECD weights are used. Since the weights are annual, we used the weight of the j say year for the four quarters of this year and the 1970 weight for the period 1962-1970. We used the 25 OECD countries, as Greece's competitors. "Historical Statistics" and "Main Economic Indicators", OECD.

Y^f = foreign income (either Y^f_{gdp} or Y^f_{ip})

Y^f_{gdp} = is the index of GDP (at constant 1990 USD prices) of five geographical areas, namely: EU(15), USA, former Yugoslavia, Middle East and Africa. The classification of Middle East countries is according to that of the U.N. The above index is derived as:

$$Y^f_{i,gdp} = \Sigma\, W_{ij}\, Y_{ij}$$

where W_{ij} = the weight which represents the share of Greek exports to the jth area in the total Greek exports to these areas in the ith quarter, Y_{ij} = GDP of the jth area in the ith quarter, i = 1, …. 152 quarters, j = 1, …. 5 areas

Annual data (the only data available) were obtained by the U.N and OECD “National Accounts” publications and the IMF “Yearbook of International Financial Statistics”. These data were benchmarked to form quarterly data using the procedure SPATQ of the TROLL computer program.

Y_{ip}^{f} = is the index of industrial production (1990=100) of OECD countries. Source: “Main Economic Indicators”, OECD.

K = is the net fixed capital stock in manufacturing derived from data on gross fixed capital formation at constant prices. Since no quarterly data are available we picked up annual figures for the period 1949-1999 of investment in machinery and transport equipment and of investment in buildings separately at 1970 constant prices, with an average service life of 20 years for the former and 50 years for the latter (following the practice of the National Statistical Service of Greece). We assumed a benchmark estimate of the capital stock at the beginning of 1948 of Dr. 13.4 billion at 1970 prices (Kintis, 1973). Since this benchmark value refers to total capital stock, we derived estimates for machinery and buildings separately (for the first year of the sample) by applying the shares of these two items in total capital stock in large-scale industry based on Kinti’s (1973) study. The above values were used to compute the net capital stock taken as benchmark estimate for the sample period. The computations were performed by using the perpetual inventory method:

$$K_t=I_{t-1}+(1-\delta)\, K_{t-1}, \qquad (1)$$

where K = capital stock, I = investment, δ = the depreciation rate (=2.5/N), N = the service life of machinery and buildings.

In expression (1) it is assumed that economic depreciation is proportional to the capital stock. Therefore, the time path of its service life is geometrically declining. The definition of δ=2.5/N implies that the depreciation rate is geometrically declining and not according to a straight line where δ=1/N. The constant 2.5 has been taken from Eisner-Nadiri (1968) and Jorgenson-Siebert (1968). By taking δ=2.5/N, the depreciation rate is assumed 0.125 and 0.05 on average per year for machinery (and transport equipment) and buildings respectively, while by taking δ=1/N the depreciation rate is 0.05 and 0.02 respectively. The annual data of the estimated capital stock were benchmarked to form quarterly data using the frequency conversion method of the EVIEWS program. For the capital stock of buildings a linear interpolation was applied, while for machinery a cubic interpolation was considered appropriate. The total capital stock series was obtained by summing the above two series.

APPENDIX B

Figures B1-B4 present recursive eigenvalues, and Figures B5-B8 show the recursively estimated LR test statistic for the overidentifying restriction (price homogeneity) for export equations (1.1)-(1.4).

Fig. B1: Eq. 1.1

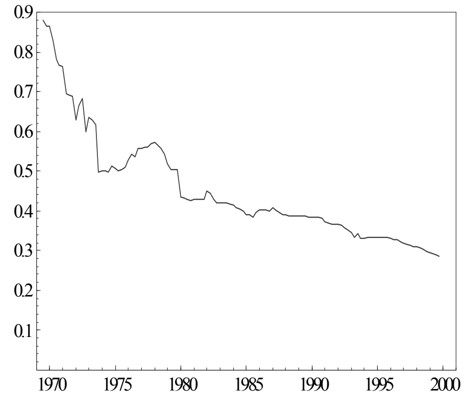


Fig. B2: Eq. 1.2

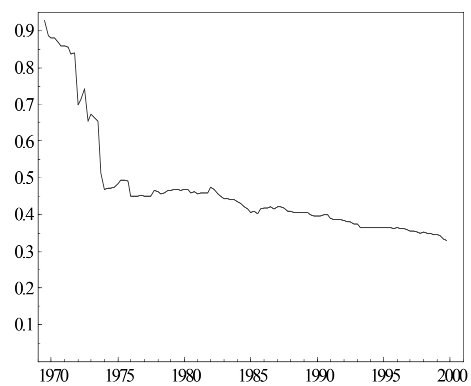


Fig. B3: Eq. 1.3

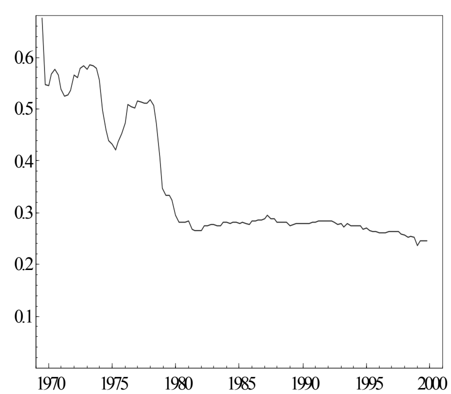


Fig. B4: Eq. 1.4

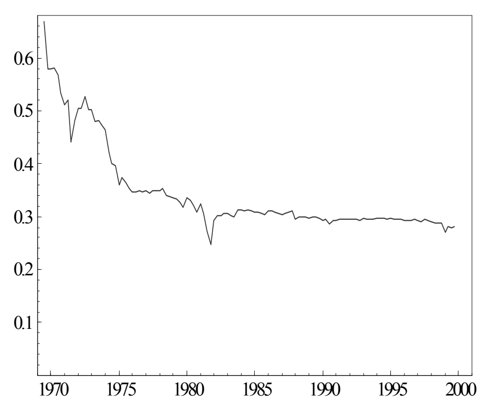


Fig. B5: Eq. 1.1

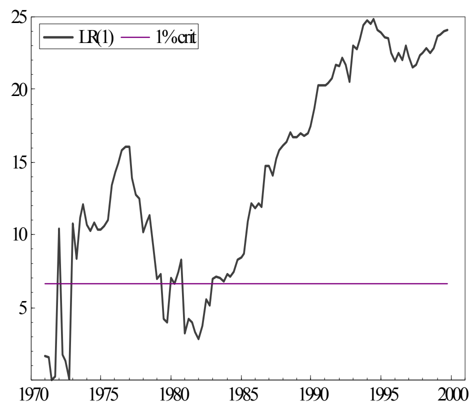


Fig. B6: Eq. 1.2

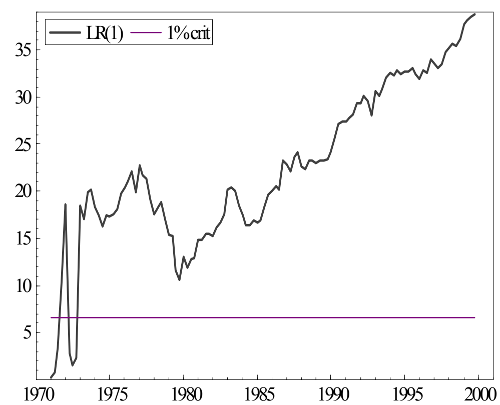


Fig. B7: Eq. 1.3

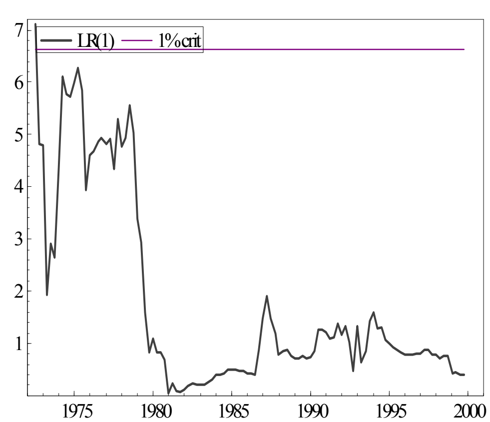
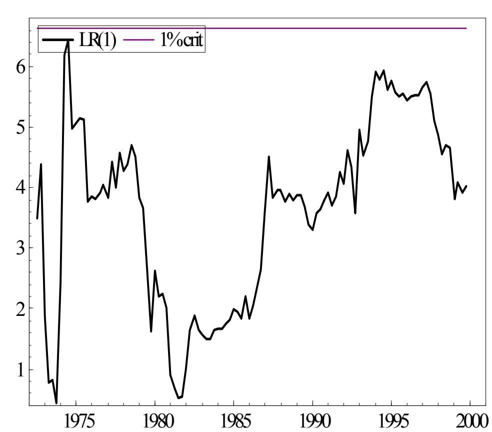


Fig. B8: Eq. 1.4



APPENDIX C

Stability of Dynamic Export Equations (1.1-1.4)

Figures C1-C8 present Quandt tests for export equations (1.1)-(1.4) and (1.1a)-(1.4a).

Fig. C1: Eq. 1.1

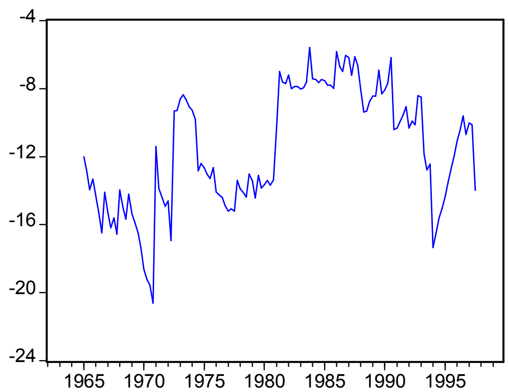


Fig. C2: Eq. 1.2

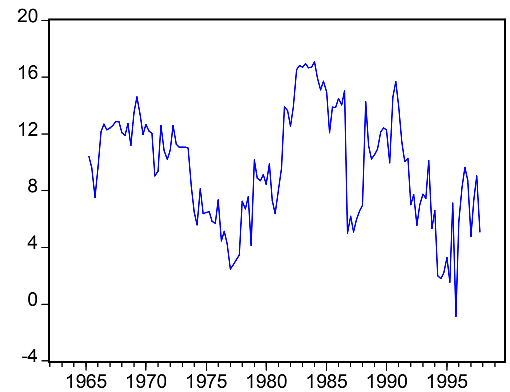


Fig. C3: Eq. 1.3

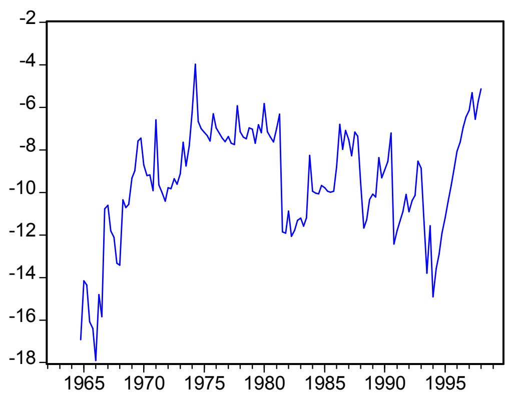


Fig. C4: Eq. 1.4

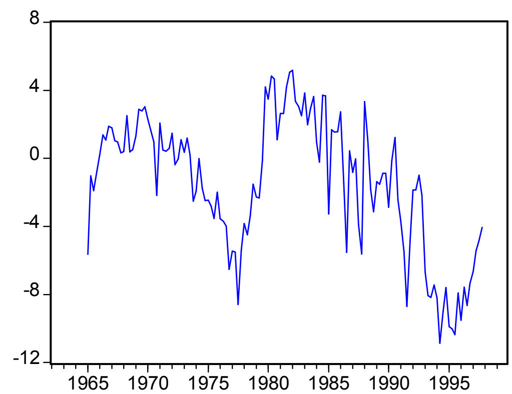


Fig. C5: Eq. 1.1a

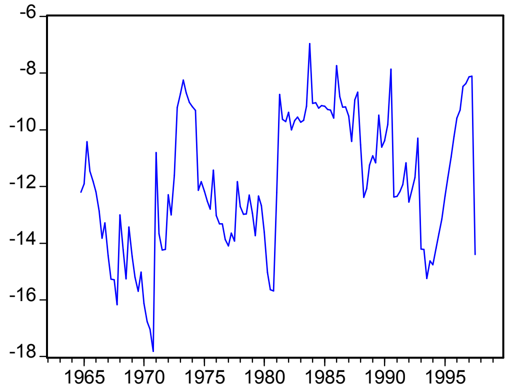


Fig. C6: Eq. 1.2a

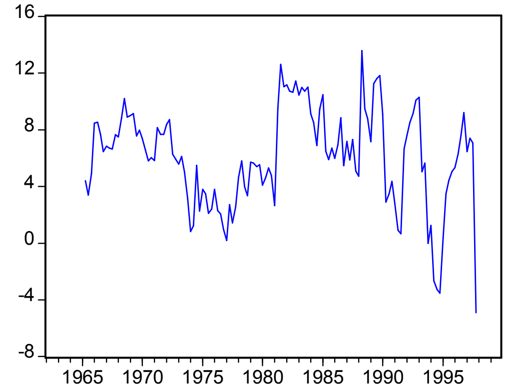


Fig. C7: Eq. 1.3a

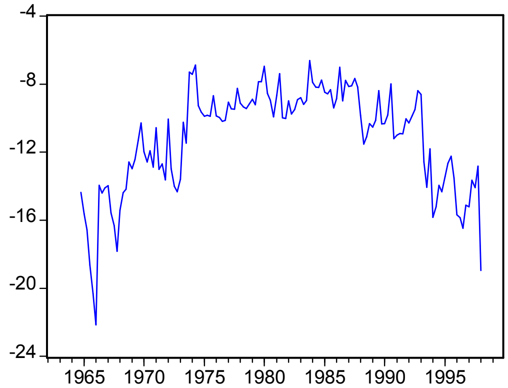
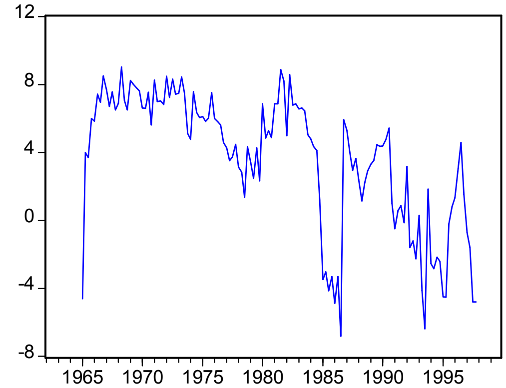


Fig. C8: Eq. 1.4a



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