AN ECONOMETRIC ANALYSIS OF THE FOREIGN TRADE OF GREECE

by

JOHN G. HALIKIAS

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Department of Economics
University of Warwick
COVENTRY CV4 7AL

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To my wife Mary and my son Gerasimos

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DECLARATION

The work contained in this thesis was the result of original research conducted by myself under the supervision of Professor Kenneth F. Wallis and all sources of information have been specifically acknowledged by means of references.

To my knowledge none of the work contained in this thesis has been previously submitted for examination.

John G. Halikias.

SUMMARY

This study is mainly concerned with the estimation of Greece's dynamic import and export demand equations for goods over the period 1954-1976.

Some recent empirical studies of the foreign trade of various countries are first presented and the approaches that have been pursued are described as a preliminary to the development of our own theoretical framework. Foreign trade elasticities are then estimated to test various hypotheses in the theory of international trade and, subsequently, to be used in the formulation of economic policy. Some applications of the empirical findings to the trade balance of Greece are attempted. In general the trade balance of Greece is found to be sensitive to both relative price changes and the growth rates of Greece and its trading partners.

In the course of the empirical work a number of methodological points of importance in applied econometrics arise. In particular we are concerned with the empirical specification of dynamic models and the resulting hypothesis-testing problems. Two model selection procedures for the empirical specification of dynamic models are described and their performance is evaluated in the context of our large scale empirical study. The first procedure begins with the simplest (static) form of the relationship and tests are performed to determine whether it is necessary to consider more general specifications. An alternative method begins with a general unrestricted dynamic model and then attempts to reduce the number of parameters needed to specify the data generation process. These two procedures are applied to every import and export demand equation we consider and the preferred specification to which each approach leads is reported. It is found that the possibility of conflict between the two procedures increases as higher-order dynamic models are considered.

The use of monthly data for the period 1954-1976 instead of annual or quarterly observations, which have been employed in previous work on models of foreign trade, constitutes a further novelty of our approach, and the resulting problems and their practical solution are described.

CHAPTER I

INTRODUCTION

1. Objectives of the Thesis

International trade has always played a vital role in a country's economic activity, and among the various branches of economics is one for which extensive and detailed statistical information exists. Given also that quantitative relationships between economic variables are invaluable tools in economic policy, many investigators have analysed these data to obtain import and export demand equations either for describing the responses of imports and exports to changes in prices and the like, or for forecasting purposes.

The present study is mainly concerned with the estimation of Greece's dynamic import and export demand equations for goods over the period 1954-1976. Foreign trade elasticities have been estimated to test various hypotheses in the theory of international trade and, subsequently, to be used in the formulation of economic policy. This study has been motivated by the lack of any dynamic model of the foreign trade of Greece, in particular, and by any recent study of its foreign trade, in general.

The present study is also concerned with a number of methodological points of importance in applied econometrics , and though the particular application chosen is that of estimation of import and export demand functions, the points have a much wider relevance. In particular we are concerned with the empirical specification of dynamic models and the testing of a series of hypotheses involved.

The lack of detailed information from economic theory on the dynamic structure of economic relationships has caused researchers to rely on empirical specification procedures for model building with economic time series data. The traditional model selection procedure is to experiment with the techniques suggested by the distributed lag literature and this approach has been followed so far, at least as far as import and export demand functions are concerned (see chapter III).

In this study we describe two alternative methods for the empirical specification of dynamic models, and we try to evaluate their performance in the context of our large scale empirical study. In particular, these two procedures have been applied for every import and export demand equation we consider and the preferred specification to which each approach leads is reported.

In the first procedure we start from the static form of the relationship without assuming anything about dynamics. Then, testing for serial correlation in the residuals enables us to specify the dynamic form of the relationship. This approach is based principally on the fact that misspecified dynamics may result in a serially correlated disturbance and considers tests which check this and allow us to discriminate between stochastic specifications and dynamic specifications.

The above method is a stepwise approach which allows us to extend the dynamic specification in a systematic way when the relevant tests indicate that the particular model estimated, at any stage of the procedure, is inadequate. That is we start with the simplest model and we test if it is necessary to consider a more general one. An alternative method is to begin with a general unrestricted dynamic model and then attempt to reduce the number of parameters needed to specify the data generation process. Under this alternative we have a maintained hypothesis which is a general unrestricted dynamic model with the maximum number of lags for all variables, and tests are carried out to ascertain whether restricted versions of it are consistent with the data.

The use of monthly data from the period 1954-1976 instead of annual or quarterly observations, which have been employed in previous work on models of foreign trade, constitutes a novelty of our approach. In the context of monthly data we describe the application characteristics of the above empirical dynamic specification procedures.

Applying the above model selection procedures, the preferred specifications and their numerical estimates of the import and export demand functions of Greece are obtained, and some applications of these estimates to the trade balance of Greece are attempted. Finally, we discuss the experience gained from the application of the above model selection procedures, and we deal briefly with the implications of the empirical findings for Greek trade policy.

2. Structure of the Thesis

In chapter II, the main characteristics of the foreign sector of the Greek Economy over the period 1954-1976 are briefly outlined, and we discuss briefly previous research on Greek imports and exports.

In chapter III, we present some recent empirical studies of the foreign trade of various countries, describing in general terms the approaches that have been pursued.

Chapter IV deals with the description of our model selection procedures, their characteristics in relation to monthly data as well as their computational aspects. Also, the basic model we adopt for the Greek foreign trade which serves as a baseline for our subsequent empirical analysis, is presented.

The presentation and discussion of the empirical findings are contained in chapters V and VI. In chapter V an attempt is made to obtain the preferred specifications and their numerical estimates of the import demand equations, while the measurement of factors determining the foreign demand for Greek exports is treated in chapter VI.

In chapter VII, we examine whether a joint estimation of the import and export demand equations could result in more efficient estimates. In particular, using residual cross correlation functions, we examine if contemporaneous or lagged correlations appear among the residuals of the various import and export demand equations.

Chapter VIII is concerned with some applications of the empirical findings to the trade balance of Greece. Specifically, predictions are made for the country's trade balance under different growth rates in Greece and its trading partners, and different relative price changes, showing at the same time the sensitivity of the country's balance of trade to these various assumptions.

Finally, in chapter IX we discuss the experience we gained from the application of the above model selection procedures, and we deal briefly with the implications of the empirical findings for Greek trade policy.

CHAPTER II

THE FOREIGN SECTOR OF THE GREEK ECONOMY

In section 1. of this chapter we describe the main characteristics of the foreign sector of the Greek economy over the period 1954 - 1976. In section 2. we discuss briefly previous research on Greek imports and exports.

1. Trends of the Foreign Sector of Greece in the Period 1954 - 1976

Greece, a small country with limited soil and subsoil resources, is highly dependent on international trade. This is mainly due to the country's present stage of development, the small size of the economy, and to the entire lack of fuels and some basic raw materials. The dependence of the country on the foreign trade increased slowly during the period 1954-1972 and more rapidly during the period 1973-1976. Thus, as it can be seen from table 1, the ratio of the sum of imports and exports to gross national income, increased from 0.252 in 1954 to 0.284 in 1972. But, during the period 1973-1976 the above ratio increased considerably and it amounted to 0.425 in 1976. Table 1 also shows that during the period 1954-1972 the export-income ratio did not vary significantly, while the corresponding import-income ratio increased from 0.172 in 1954 to 0.207 in 1972. But during the period 1973-1976 the export-income ratio increased somewhat faster than the import-income ratio. Conse-

TABLE 1

RATIOS BETWEEN EXPORTS, IMPORTS AND GROSS NATIONAL INCOME

(In million Drachmae at current prices)

	Imports (c.i.f.)	Exports (f.o.b.)	Trade Balance	Gross National Income	Imports+Exports / Income	Imports / Income	Exports / Income	Exports / Imports
1954	9901	4556	- 5345	57467	0.252	0.172	0.080	0.460
1958	16946	6953	- 9993	85162	0.281	0.199	0.082	0.410
1960	21060	6096	- 14964	95174	0.285	0.221	0.064	0.289
1965	34012	9833	- 24179	161586	0.271	0.210	0.061	0.289
1970	58750	19276	- 39474	263503	0.296	0.223	0.073	0.328
1972	70373	26125	- 44248	339554	0.284	0.207	0.077	0.371
1973	102979	42812	- 60167	441301	0.330	0.233	0.097	0.416
1974	132181	60890	- 71291	530081	0.364	0.249	0.115	0.461
1975	172041	74441	- 97600	612388	0.402	0.281	0.121	0.433
1976	221821	93812	-128009	742436	0.425	0.299	0.126	0.423

Sources: National Statistical Service of Greece, Monthly Bulletin of External Trade Statistics; National Accounts of Greece, 1948 - 1975 and 1970 - 1976.

quently, the export-import ratio declined from 0.460 in 1954 to 0.371 in 1972 whereas it increased to 0.423 in 1976.

With respect to the foreign trade of Greece, the post-war period can be divided in three sub-periods. The first sub-period includes the years 1948-1952, when unusual conditions prevailed in the economy of Greece. War damages (the civil war lasted until late 1949) and hyperinflation had severely dislocated the economic activity of the country, bringing its production to extremely low levels. At the same time severe import restrictions were imposed by the authorities to control the large deficits in the balance of payments. In 1953 Greece devalued its currency (drachma) by 50 percent in order to eliminate the fundamental disparities between domestic and international prices. Simultaneously, the authorities liberalized imports to put foreign trade on a sounder basis conducive to the economic development of the country. Since then, Greece has pursued a liberal import policy to the extend that only a small proportion of imports requires import licences. In view of the above discussion the period 1948-1953 has been excluded from our empirical analysis (see also chapter V, section 1., below).

From 1954 to 1972, we have the second sub-period, during which, the gradual reinstatement of the monetary stability of the country, resulted in the formation of better conditions in the foreign sector of the economy. Finally, the third sub-period starts from 1973, when the international price increases of raw materials and particularly of crude oil, had an unfavorable effect on the foreign trade of Greece.

TABLE 2

UNIT VALUE INDEX AND QUANTUM INDEX OF IMPORTS AND EXPORTS

(1970 = 100)

	UNIT VALUE INDEX		TERMS OF TRADE	QUANTUM INDEX	
	Imports	Exports		Imports[1]	Exports
1954	98.0	85.7	87.4	21.3	27.6
1958	96.4	96.7	100.3	35.1	37.3
1960	93.1	86.5	92.9	35.6	36.6
1965	94.7	98.2	103.7	71.4	52.0
1970	100.0	100.0	100.0	100.0	100.0
1972	112.6	104.1	92.5	122.7	130.2
1973	134.6	136.0	101.0	154.4	163.3
1974	195.4	177.0	90.6	141.0	179.3
1975	233.1	196.6	84.3	137.7	195.7
1976	259.8	215.5	82.9	148.4	225.9

[1] Excluding ships

Sources: National Statistical Service of Greece, Annual Volumes of External Trade Statistics

Due to the relatively worldwide monetary stability during the period 1954-1972, the unit value index of imports increased at an average annual rate of 1.75 percent and the unit value index of exports at an average rate of 2.46 percent. From a quantity point of view, imports increased six times and exports increased almost five times (table 2). In view of the above events, the trade deficit of the country increased from 5,345 million drachmas in 1954 to 44,248 million drachmas in 1972, namely eight times more, whereas, as already mentioned, the export-import ratio declined from 0.460 in 1954 to 0.371 in 1972 (table 1).

Over the same period, some noticeable structural changes of the composition of the external trade of Greece, took place. As can be seen from table 3, which shows the composition of imports and exports according to the Standard International Trade Classification (S.I.T.C.), the share of food and manufactures in total imports decreased, whereas the share of imported machinery and transport equipment was doubled. With respect to exports, the share of the traditionally exported agricultural commodities (tobacco, raisins and oil) in total exports, declined from 65.1% in 1954 to 18.0% in 1972. On the contrary, the share of more dynamic agricultural products, such as fresh fruit, vegetables and cotton, in total exports, increased from 8% in 1954 to 20% in 1972. A noticeable increase has been also noticed in the participation of exported chemicals and manufactures, whose shares in total exports increased from 3.4 and 5.8 percent respectively, in 1954 to 7.4 and 32.6 percent in 1972. Thus, Greek exports do not depend any more on a few

TABLE 3

PERCENTAGE DISTRIBUTION OF IMPORTS AND EXPORTS

SITC Sections	Imports[1]						Exports					
	1954	1960	1970	1972	1975	1976	1954	1960	1970	1972	1975	1976
0. Food	16.8	14.6	12.2	10.9	10.2	9.4	24.9	25.6	22.9	25.0	22.3	22.5
1. Beverages and tobacco	-	-	0.2	0.2	0.1	0.2	42.8	37.1	17.5	16.0	8.2	8.6
2. Raw materials	14.9	12.7	10.6	9.5	9.0	8.5	15.7	25.2	16.9	14.1	8.8	10.0
3. Mineral fuels, lubricants etc.	14.0	10.2	8.7	10.8	25.4	25.2	-	-	1.0	1.2	11.0	5.8
4. Animal and vegetable oils and fats	0.1	0.2	0.8	0.2	0.4	0.3	6.7	2.1	0.8	1.5	1.8	0.8
5. Chemicals	9.3	10.6	10.2	10.7	10.0	10.2	3.4	4.1	7.2	7.4	5.8	4.0
6. Manufactures	23.8	24.1	19.7	18.7	15.9	15.3	5.4	4.1	28.6	26.4	28.7	31.7
7. Machinery and transport equipment[1]	17.4	24.2	33.9	35.4	26.1	27.7	0.7	0.9	1.5	2.2	3.9	4.9
8+9. Miscellaneous manufactures	3.7	3.4	3.7	3.6	2.9	3.2	0.4	0.9	3.6	6.2	9.5	11.7
(0-9) TOTAL	100.0	100.0	100.0	100.0	100.0	100.0	100.0	100.0	100.0	100.0	100.0	100.0

[1] Excluding ships

Sources: Elaboration of External Trade Statistics of the National Statistical Service of Greece.

agricultural products whose supply and quality depend on each time weather's conditions.

During the last period 1973-1976 the price increases of the imported and exported commodities were very high (annual rate of increase 24.5 percent and 16.6 percent respectively - table 2) as a result of the considerable price increases of crude oil and raw materials specifically, and more generally because of the international and domestic inflationary pressures. Moreover, the imported inflation was reinforced because of the devaluation of the Greek currency (drachma) against the U.S.A. dollar and the main currencies of Western Europe, which took place in 1975.

In terms of quantity, imports increased initially in 1973, but then declined in 1974 and 1975 as a result of the reduced economic activity caused by the Cyprus crisis during that period; finally increased again in 1976. On the contrary, during the same period, exports increased at an average annual rate of 11.4 percent, and though the trade deficit increased more, the export-import ratio was slightly improved and it increased from 0.416 in 1973 to 0.423 in 1976.

The deficit of the balance of trade was covered for the most part by the increasing receipts from sales of services (tourism and shipping). In particular, during the period 1954-1972 the above receipts increased thirteen times, financing about 27 percent of the merchandise imports in 1972, whereas in 1954 they covered only 17 percent. These receipts increased more during the period 1973-1976, but they were affected from the non-economic events of 1974 (Greek-Turkish disputes) which

TABLE 4

BALANCE O PAYMENTS

(In million Drachmae at current prices)

	1954	1958	1960	1965	1970	1972	1974	1975	1976
A' CURRENT TRANSACTIONS									
1. Exports - goods (fob)	4803	7216	6242	9960	19378	26203	62098	74787	93948
2. Less: Imports - goods (cif)[1]	8541	16092	15840	33113	49262	67115	130724	149447	176869
I. Balance of trade	-3738	-8876	-9598	-23153	-29884	-40912	-68626	-74660	-82921
3. Sales of non-income services	375	850	949	2182	3576	4876	8580	12905	13318
4. Expenditure of non-residents	1049	1914	2408	3980	7034	13221	17187	22123	32275
Less:									
5. Purchases of non-income services	1450	586	823	1416	2850	4452	8358	24430	28612
6. Expenditure of residents abroad	465	676	899	1947	2884	4156	5613	6685	7553
II. Balance of services	- 491	1502	1635	2799	4876	9489	11796	3913	9428
7. Income payments from the rest of the world	975	1601	2213	4388	7777	13027	24744	26761	34906
8. Less: Income payments to the rest of the world	102	142	264	758	2274	3450	6876	7554	9964
III. Net income from the rest of the world	873	1459	1949	3630	5503	9577	17868	19207	24942
IV. Balance of goods, services and incomes (I + II + III)	-3356	-5915	-6014	-16724	-19505	-21846	-38962	-51540	-48551
V. Current transfers from the rest of the world (net balance)	1406	2681	3006	6241	10203	17126	20325	23342	27402
VI. Balance of current transactions (IV + V)	-1950	-3234	-3008	-10483	- 9302	- 4720	-18637	-28198	-21149
B' CAPITAL TRANSACTIONS									
1. Capital transfers from the rest of the world (net balance)	1739	912	1459	469	54	21	30	431	333
2. Net lending	211	2322	1549	10014	9248	4699	18607	27767	20816
TOTAL (1 + 2)	1950	3234	3008	10483	9302	4720	18637	28198	21149

[1] Excluding ships operating overseas

Source: National Accounts of Greece, 1958 - 1975, and 1970 - 1976 had an unfavorable effect on tourism. Also, foreign exchange earnings from transportation were affected from the depression in the international transportation market, which occured during the years 1975-1976. In view of the above, the receipts from sales of services financed only 19.7 percent of merchandise imports in 1974, 23.4 percent in 1975 and 25.8 percent in 1976.

Table 4 shows the balance of payments of Greece, during the period 1954-1976, according to the operating classification of National Accounts. As can be seen from the above table, the deficit of the balance of goods, services and incomes, increased from 3,356 million drachmas in 1954 to 51,540 million drachmas in 1975, but decreased to 48,551 mil. dr. in 1976. Till 1951, the majority (about 90%) of the above deficit was covered by capital transfers (U.S.A. aid and reparations).

Since 1952 the U.S.A. aid was decreasing gradually and the deficit of the balance of goods, services and incomes was covered by the current transfers from the rest of the world (mainly emigrant remittances and workers' earnings from Europe) and lending. Thus, the percentage of the above deficit which is covered by the current transfers increased from 6 percent in 1950 to 41.9 percent in 1954 and 52.3 percent in 1970. In 1975 it declined to 44.2% since the increased unemployment in West Europe through its effect on the workers' earnings affected unfavorably the net balance of the current transfers. However, the above ratio increased again to 56.4 percent in 1976 (table 5).

Finally, the net lending of the country increased from

TABLE 5

COVER OF THE DEFICIT OF THE BALANCE OF

GOODS, SERVICES AND INCOMES

(PERCENTAGE DISTRIBUTION)

	1950	1954	1960	1970	1975	1976
1. Current transfers from the rest of the world (net balance)	6.0	41.9	50.0	52.3	45.3	56.4
2. Capital transfers (U. S.A. aid - reparations)	91.5	51.8	24.3	0.3	0.8	0.7
3. Net lending	2.5	6.3	25.7	47.4	53.9	42.9
Deficit of the balance of goods and services	100.0	100.0	100.0	100.0	100.0	100.0

Source: National Accounts of Greece

211 million drachmas in 1954 to 20,816 million drachmas in 1976, covering in average the 47.8 percent of the deficit of the balance of goods, services and incomes, during the last four years.

2. Previous Studies of the Foreign Trade of Greece

Empirical studies of the foreign sector of the Greek economy have been carried out by Suits (1965), Adelman and Chenery (1966), Pavlopoulos (1966), Paraskevopoulos (1970), Hitiris (1972)[1], Sarantides (1973), and Prodromidis (1974). These studies vary according to the sample period, the sectoral breakdown and the explanatory variables used. All of them deal with the estimation of import demand equations, which we discuss first, whereas little attention has been paid to exports (Adelman and Chenery (1966), Paraskevopoulos (1970), Hitiris (1972) and Prodromidis (1974)).

2.1. Imports

Imports are regarded as the difference between two variables, total consumption and domestic production, and the import function is the difference between the functions explaining these two variables, that is demand and supply functions. This implies that the imported and home produced commodities are identical. But, if an imported commodity is not produced at home, or the imported and home produced commodities are not identical, the import demand coincides with the home demand for that commodity. Therefore, due to the fact that the majority of commodities imported into Greece are not produced at home, all the above investigators formulate import demand equations

[1] For a brief description of his work see also Hitiris (1968)

which simply express the country's home demand for such commodities.

The above studies cover the following periods: Suits (1953-1961), Adelman and Chenery (1950-1961), Pavlopoulos (1949-1959), Paraskevopoulos (1954-1966), Hitiris (1955-1964), Sarantides (1953-1964) and Prodromidis (1961-1969). All the authors estimate linear regression equations using annual data for the periods we just mentioned. Paraskevopoulos estimates also double logarithmic forms using quarterly data, whereas Hitiris (1972) uses only quarterly data for the estimation of simple linear form equations.

The level of aggregation at which these empirical studies have been carried out, varies from author to author, and there is no sound explanation for that. However, the availability of data at the time these studies were carried out affected the sectoral breakdown followed by each author. The National Statistical Service of Greece started the compilation of foreign trade indices (quantum and unit value at the one-digit SITC level of aggregation) at the end of 1956. They cover the period 1951-1953 with annual data and the period 1954 onwards with monthly and annual data, the first figures being available by the end of 1958. Until then only quantities and values, in current and constant prices, were provided by the National Statistical Service and the Ministry of Coordination (National Accounts). This has also affected, as we shall see below, the explanatory variables included in import demand equations. The table below gives the major groups of commodities for which import demand equations were estimated by the studies in question.

TABLE 6

LEVEL OF AGGREGATION ADOPTED BY PREVIOUS STUDIES ON GREEK IMPORT DEMAND EQUATIONS

Suits (1953-1961)[a]	Adelman and Chenery (1950-1961)[a]	Pavlopoulos (1949-1959)[a]	Hitiris - Paraskevopoulos (1955-64)[b] (1954-66)[a, b]	Sarantides (1953-1964)[a]
- Agricultural commodities (i) animal products and fish (ii) luxury agricultural products (sugar, coffee, cocoa etc.) (iii) plant products with domestically produced substitutes (cereal grains and other plant crops) (iv) edible oils - Manufactures (i) private consumer goods (ii) private non-consumer goods	- Foods, beverages and animal and vegetable oils - Crude materials plus mineral fuels and chemicals - Manufactures	- Raw materi-als - Consumption goods - Investment goods	- Foods (SITC 0) - Crude materials (SITC 2) - Fuels (SITC 3) - Chemicals (SITC 5) - Manufactures (SITC 6) - Machinery and Transport equipment (SITC 7)	- Foods - Raw materials, fuel, and intermediates - Manufactures - Construction materials (iron, steel, timber, copper etc.) - Capital goods and construction materials - Machinery - Transport equipment

[a]Annual data

[b]Quarterly data

Except Suits, all the authors have also estimated demand equations for total imports, whereas Paraskevopoulos has also estimated import demand equations for individual commodities as well, such as meat, fish, dairy products, coffee, cocoa, sugar, passenger cars, textile products without raw wool, raw wool and services. Prodromidis (1974) examines import according to the purposes of a sectoral planning model of the Greek economy which is under preparation at the Athens Center of Planning and Economic Research. He classifies the importable commodities in accordance with the input-output classification system of Greece and not according to their use as is usual. Even in this most recent one work the construction of new series of data is necessary and the study therefore covers only the period 1961-1969[2].

According to the formulation of the Greek import demand functions by the above investigators, the main explanatory variables which have been included in their equations are activity variables and relative prices. As was mentioned before, unit value indices are listed only for the S.I.T.C. groups of commodities and therefore only Paraskevopoulos and Hitiris include import prices in their disaggregated import demand equations, deflated by the indices of domestic prices. In Suits' import demand equations for manufactured consumer goods, animal products and fish, and luxury agricultural products, the import prices are implicit deflators obtained as the ratio of the current va-

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2 For a brief description of his work see also Prodromidis (1975)

lue of imports to value at constant prices. Adelman and Chenery employ the same relative price of imports in all equations. This is the price index of total imports divided by the GNP deflator. The rest of the authors use only income or activity variables to explain the behaviour pattern of imports (Pavlopoulos (1966), Sarantides (1972)).

The explanatory variables used to measure the economic activity of Greece, differ from author to author depending upon each investigator's classification. Income, in various forms, has been included in all the models; i.e. disposable income, GNP, net national income. In Greece quarterly figures for income are not listed and in both Paraskevopoulos' (1970) and Hitiris' (1972) works the index of industrial production has been chosen as an income-proxy. Value added by manufacturing activity has been employed as explanatory variable in import demand equations for manufactured non-consumer goods, raw materials, fuels and intermediates (Suits (1965), Pavlopoulos (1966) and Sarantides (1972)). Also, various forms of investment activity and expenditure (gross fixed capital formation, investment in housing, building and other construction, machinery, transport equipment etc.) have been used to explain the imports of investment goods, construction materials, capital goods, and machinery and transport equipment (Pavlopoulos (1966) and Sarantides (1972)).

Suits (1965) has also included in his import demand equations for plant products and edible oils, the lagged stock of cereals and the lagged stock of oil, respectively, to take account of the dynamic effects of past purchases. Finally, only

Pavlopoulos (1966) has attempted to include in his equations one-period lagged imports, but only for the import demand function for consumption goods he obtained significant results.

2.2. Exports

Foreign demand functions for Greek exports have been formulated in a way analogous to that used for the country's import demand relationships, and have been expressed as functions of world activity or income variables and the ratio of Greek export prices to export prices of Greece's competitors.

However, few empirical studies have been carried out on Greek exports because during the periods covered by the above studies, Greek exporting activity was low and the major part of exports consisted of a few agricultural products (Adelman and Chenery (1966), Paraskevopoulos (1970), Hitiris (1972), and Prodromidis (1974)). The table below shows the groups of commodities for which export demand equations have been estimated by the above investigators.

In Adelman and Chenery's work each group of commodities is expressed simply as a linear function of time. Paraskevopoulos ((1970), ch. 6) and Hitiris ((1972), ch. 3) estimate linear (and double logarithmic - Paraskevopoulos) regression equations with annual data which refer to the periods 1951-1966 and 1954-1966 respectively. Their explanatory variables are, as mentioned before, relative prices and economic or activity variables of the areas of destination (volume of food consumption, cigarette consumption, cotton production, index of industrial production,

TABLE 7

LEVEL OF AGGREGATION ADOPTED BY PREVIOUS STUDIES ON GREEK EXPORT DEMAND EQUATIONS

Adelman and Chenery (1966)	Paraskevopoulos (1970)	Hitiris (1972)
-Food, beverages & tobacco, and animal & vegetable oils	-Food	-Food and live animals (SITC 0)
-Crude materials, mineral fuels and chemicals	-Dried fruits (currants, raisins and dried figs)	-Beverages and Tobacco (SITC 1)
-Manufactured goods, and machines and transport equipment	-Tobacco	-Inedible crude materials (SITC 2)
-Services	-Cotton	-Chemicals (SITC 5)
	-Non-cotton raw materials	-Manufactures (SITC 6)
	-Manufactures and chemicals	
	-Services	

GNP and population). They have also estimated export demand equations for the total of Greek exports. Finally, Prodromidis (1974, ch. 5) estimates linear and double logarithmic export demand functions over the period 1961-1969 for groups of commodities according to the input-output classification system of the Greek economy.

Except for Adelman and Chenery (1966) who experimented also with two stage least squares and limited information methods, all investigators used the ordinary least squares method for the estimation of their import (and export) demand equations. They adopt the hypothesis that since Greece is a small country with a small participation in international trade import prices can be treated as predetermined variables in the import demand equations. An analogous treatment of the income of Greece prevails in the above studies based on the assumption that the demand for these commodities has a negligible effect on income or economic activity of Greece. Therefore they presume that the disturbances are independent of the explanatory variables and so least squares estimates are free of simultaneity bias. Over the periods covered by the above studies, the Greek economy was characterized by a rapid rate of growth. The result is a high intercorrelation among the various time series and most of the authors faced the problem of the deterioration of statistical significance of regression parameters when they attempted to include additional explanatory variables (Suits (1965), Pavlopoulos (1966), Sarantides (1972) and Prodromidis (1974)).

Finally we note that all the estimated import and export demand equations are static in that all variables relate to a single time period. The use of annual data and the relatively short sample periods may explain the absence of lagged variables and the problem of autocorrelation becomes less serious. Only in Hitiris (1972) and Sarantides (1972) do any equations exhibit serial correlation in the disturbances.

Thus there is little recent work on the foreign trade of Greece. However taking into consideration that during the last decade industrialization in Greece has expanded and its exports of manufactures have increased, plus the fact that the available statistical data have increased, we believe that possibilities for more fruitful modelling of Greek foreign trade now exist.

CHAPTER III

GENERAL ISSUES IN MODELS OF FOREIGN TRADE

In this chapter we briefly consider the main empirical studies of the foreign sectors of various countries carried out during the last decade. Our intention is to discuss the methodology adopted and the hypotheses selected for testing, elucidating any generalizations about these approaches that emerge. For convenience we shall refer first to the estimated import demand equations and then to the export demand equations (for a more extended survey see Stern at al (1976)).

1. Import Demand Equations

The demand for an imported commodity reflects the difference between the home demand for and the home supply of that particular commodity. This implies that the imported and home produced commodities are identical (or homogeneous). Then the demand for imports is treated as the excess demand function of traditional international trade theory. But, if the imported commodities are not produced at home, or the imported and home produced commodities are not identical, as happens with most of them, then import demand coincides with home demand and the problem of estimating the import demand equation for a group of commodities is generally reduced to the problem of estimating the country's home demand equation for such commodities. The majority of investigators adopt the latter hypothesis and they formulate import demand functions which determine imports, usually in terms of activity variables, relative prices and other variables reflecting each country's prevailing conditions as we shall see.

However the concept of identical commodities is a relative matter depending upon one's classification. Whether a commodity is entirely foreign-produced or also the subject of home production depends on the level of aggregation in the commodity statistics. The empirical studies to which we refer have been carried out at a level of aggregation which varies from author to author. We note that Whitley (1977) classifies U.K. imports into two major groups of commodities, namely manufactures and semi-manufactures, whereas Kreinin (1973) works at a disaggregate level of 56 groups of commodities for U.S. imports. In most of these studies a single aggregate import demand equation is not estimated, but the elasticities of total imports with respect to income and prices are derived from the elasticities of the disaggregated functions weighted by the shares of each import category in total imports. Thus aggregation bias may be reduced: see Barker (1970) for a detailed analysis of its sources.

Economic theory offers little guidance on the appropriate functional form for import demand relationships and it rests upon each investigator to decide what form is more convenient for purposes of estimation. The linear import demand function implies a constant marginal propensity to import and declining elasticities with respect to the explanatory variables as imports increase, and has been adopted by a limited number of investigators (Ball and Marwah (1962), Turnovsky (1968), Dutta (1964), Kwack (1972), Rhomberg and Boissonneault (1965)). On the other hand the majority of investigators adopt the double-logarithmic form for import demand relationships which yields direct estimates of the relative price and income elasticities and assumes constant elasticities, which seems more plausible. However, this specification constrains the import demand elasticity with respect to import prices to be equal in magnitude but opposite in sign to the elasticity with respect to domestic prices. In a test carried out by Murray and Ginman (1976) with an aggregate import demand equation for Canada, the above specification was rejected.

The final form of the import demand equation to be estimated is developed in various ways. Most of the studies simply write down the relationship between actual import demand and its determinants according to the relevant hypothesis. Import demand functions are also developed from the combination of the relationship between long-run (or desired) demand for an imported commodity and its determinants and a partial adjustment mechanism (see Aurikko (1975), Goldstein and Khan (1976), and Yadav (1975)). Whitley (1977) also includes the response of imports to discrepancies between desired and actual stocks, obtaining finally an import demand equation whose dependent variable is not actual imports, as all other cases, but the ratio of current imports to their value lagged one period. Such procedures yield dynamic import demand functions, with lagged imports included as predetermined variables, but, as we shall see below, other forms of dynamic relationship have also been estimated.

The majority of the investigators use the traditional single-equation methods to estimate import demand functions, by regressing actual imports on explanatory variables, such as import prices and domestic income. However, attempts have also been made to develop equations which describe the behaviour of import supply and import prices. Turnovsky (1968) builds a three-equation model determining import quantities and prices. The first equation regresses import demand on income, relative prices, net overseas assets and the country's exports. The second equation determines import supply in terms of overseas assets, exports and a weighted activity variable of the country's import-supplying countries (see p. 775), while the third equation states the equality between import demand and supply.

Ahluwalia and Hernández-Catá (1975) give a system of two equations in which the first expresses imports as a function of income and distributed lags of import and domestic prices, while the second, following a profit-maximization theory, determines the level of import prices in terms of distributed lags of foreign market prices, exchange rates, domestic prices and foreign capacity utilization variables. The latter approach attempts to fill the existing gap in the literature which traditionally treated import price as exogenous variables; that more attention should be paid to this had been pointed out by Prais (1962, p. 577).

Now we turn to discuss the explanatory variables which have been used as determinants of import demand. Since the import demand equations have been generally treated as home demand equations, as mentioned above, import prices and domestic prices, usually in the form of relative prices, and domestic income or an activity variable have been included in all the models. Moreover, some individual attempts have been made to include other variables which are expected to explain a substantial part of the variation of imports: for example, stocks have been included to take account of the dynamic effects of past purchases (Kwack (1972), Rhomberg and Boissonneault (1965), Whitley (1977), Hibberd (1977), Rees and Layard (1971), and Hibberd and Wren-Lewis (1978)).

Khan and Ross (1975) try to separate import demand into its cyclical and secular components. They include in their import demand function, in addition to other explanatory variables, real domestic income and its trend level (defined as potential income): a series of its values for a number of countries is provided by O.E.C.D. An analogous approach is adopted by Barker (1977) who includes as an explanatory variable in his import demand equation the ratio of the total demand of the commodity under consideration to the trend component of the total demand.

Another explanatory variable which is used by investigators dealing with British imports, is a capacity utilization variable, expressed as a function of the proportion of manufacturers working at full capacity (Barker (1976), Whitley (1977), Hibberd (1977), Rees and Layard (1971), and Hibberd and Wren-Lewis (1978)). This variable is introduced to take account of temporary increases in imports due to shortages of domestic supply. Also time trend is introduced in cases where a movement in imports cannot be accounted for in terms of changes in activity or relative prices, but is the result of changes in tastes, technology and the like (Barker (1976 and 1977), Dutta (1964), Hibberd (1977), Rees and Layard (1971)).

Only a small number of recent investigators adopt static relationships in which all variables relate to a single time period (Ball and Marwah (1962), Kwack (1972), Khan and Ross (1975), Kreinin (1973)). In these cases, where the hypothesis to be tested is that consumers adjust themselves to changed conditions within this single period, short-run and long-run elasticities are assumed to be equal. But the majority of the models include lagged variables since it is more realistic to assume that the responses of imports to changes in the explanatory variables can be delayed as purchasers take time to readjust their spending patterns. Whenever one-period lagged imports are included in the import demand equation as predetermined variable, long-run (or equilibrium) elasticities can also be derived (Turnovsky (1968), Dutta (1964), Rhomberg and Boissonneault (1965), Goldstein and Khan (1976), Yadav (1975), Whitley (1977), and Houthaker and Magee (1969)).

Relative prices enter Barker's (1970, 1976, and 1977) estimated import demand equations with a single period lag, whereas in Price and Thornblade's (1972) model, relative prices follow a distributed lag pattern over two periods. Also Almon distributed lags have been adopted to specify the time shape of the reaction of demand to changes in relative prices and activity variables (Aurikko (1975), Ahluwalia and Hernández-Catá (1975), Hibberd (1977), and Rees and Layard (1971)).

Finally, all the authors have estimated their models by ordinary least squares. Only Barker (1977) applied non-linear least squares to take account of first order serial correlation in the disturbances; Whitley (1977) estimated his equations by the instrumental variables method, treating the activity variable and lagged imports as endogenous and using as instruments current and lagged values of exports, public authorities' fixed investment and current expenditure, lagged relative prices and the level of stocks lagged by two quarters.

2. Export Demand Equations

Exports have also attracted investigators' interest since the export sector plays a central role in a country's economy both in terms of generating employment and by providing the means to pay for imports. The majority of the investigators formulate export demand equations in a way analogous to that used for import demand relationships. Particularly they develop export demand functions which determine exports in terms of world demand or activity variables, relative prices and other specific variables which are expected to contribute to the explanation of each country's exports, as we shall see below. It should be mentioned, however, that the above formulation assumes that there are no supply constraints and therefore what is specified is an export demand equation.

The majority of the researchers, as in the case of import demand functions, adopt the double-logarithmic form, which yields direct estimates of the elasticities with respect to income and prices. On the contrary only a limited number of investigators adopt the linear export demand function which implies a constant marginal propensity to export and increasing or decreasing elasticities as exports increase (Turnovsky (1968), Dutta (1964), Kwack (1972), Rhomberg and Boissonneault (1965)).

Typically, a country's export volume is explained in terms of world demand, the country's competitiveness, the profitability of exporting relative to selling at home and the pressure of internal and external demand. However, some researchers follow an indirect way to define the final form of the export demand functions, taking into consideration other factors which may affect the determination of exports. For instance, Hutton and Minford (1975) employ a mixture of demand and supply functions in the description of export sales. They consider the specification of demand and supply schedules and the behaviour of deliveries under conditions of disequilibrium between demand and supply. Supply influences enter their model in a constraining manner only when the demand for exportables from home and foreign buyers is greater than or equal to the available capacity. They also argue that the structural parameters of the export demand equation vary according to whether there is world excess demand or supply of exportables and therefore they split their estimation period according to an index of capacity utilization and estimate separate equations for each. Batchelor (1977a) introduces an econometric model of export sales behaviour, the parameters of which vary with the level of internal demand pressure measured by capacity utilization. Also in another paper on U.K. exports (1977b) he develops a model to test whether the estimated elasticities are constant over the whole sample period or whether they have changed with the changing exchange rate regime.

The competitiveness of a country's exports is a major determinant and has been taken into consideration by all the investigators. It is usually measured as the ratio of the export price to a weighted index of competitors' export prices (Kwack (1972), Rhomberg and Boissonneault (1965), Aurikko (1975), Houthaker and Magee (1969), Hutton and Minford (1975), Batchelor (1977a), Laury and Warburton (1977), Richardson (1977)). Richardson (1977) employs also the ratio of U.K. unit manufacturing cost to competitors' unit manufacturing cost, but due to lack of data he is confined only to the unit labour cost. Also all the investigators have included in their models a variable which measures the world demand for exports. This variable is either a measure of world real income (Turnovsky (1968), Kwack (1972), Houthaker and Magee (1969)) or an activity variable which measures the world's production (Dutta (1964), Winters (1976 and 1977)) or world imports or exports (Rhomberg and Boissonneault (1965), Aurikko (1975), Hutton and Minford (1975), Batchelor (1977a and 1977b), Laury and Warburton (1977), Richardson (1977)) usually in the form of a weighted index. Winters (1976 and 1977) uses also as a measure of demand for the less developed countries their capacity to import which is based on their availability of foreign exchange.

The profitability of exporting relative to selling at home has been included in export demand equations only by two authors, both dealing with U.K. exports (Winters (1976), Laury and Warburton (1977)) and is measured by the ratio of export prices to domestic wholesale prices. The effect of the internal pressure of demand on exports is examined by all the investigators who analyse British exports, because they believe that as home demand for a commodity rises, export supply is reduced. Various variables have been used to measure the internal pressure of demand, such as the domestic demand for exportables (Hutton and Minford (1975)), the index of capacity utilisation of manufacturing industry (Batchelor (1977a)), the pressure on capacity to produce which is measured by the ratio of production output to its log-trend (Winters (1976 and 1977)), the ratio of home demand to its estimated trend level (Laury and Warburton (1977)) and the index of export weighted capacity utilisation (Richardson (1977)).

In an analogous way few authors consider the effect of external demand. In Hutton and Minford's model (1975) the variable included for that reason is the world business cycle. This is a weighted index of the log-deviations of the industrial productions of the main industrial countries from their trend levels. Batchelor (1977a) uses as a variable the proportion of firms experiencing excess foreign demand, whereas Laury and Warburton (1977) measure the external demand pressure using the industrial production of the O.E.C.D. countries.

Finally, some researchers have included a time trend in their models, since exports often exhibit trends that cannot be explained by means of the above independent variables (Dutta (1964), Hutton and Minford (1975), Winters (1976 and 1977)).

The majority of the investigators analyse exports taking export prices as exogenous variables. However some attempts have been made, by authors who deal with U.K. exports, to develop equations which determine the behaviour of export prices

(Hutton and Minford (1975), Batchelor (1977a and 1977b), Winters (1976)). Export prices have been explained mainly as a function of domestic prices and competitors' export prices. The role of the home price is to measure the opportunity cost of a unit of produce exported and also to measure the actual cost of production (see Winters (1976), p. 133). The level of competitors' export prices is included in the export price equations for a reason analogous to that mentioned before in relation with the export demand functions since in most cases different exporters supply the same market. Also pressure on productive capacity and home sales have been included in the equations as a measure of domestic demand pressure to allow for the effect on prices of economies or diseconomies of scale (Batchelor (1977a), Winters (1976)).

Only a small number of recent investigators adopt static relationships is which all variables relate to a single time period (Dutta (1964), Kwack (1972), Houthaker and Magee (1969), Winters (1976)). Most authors develop dynamic relationships since it is more realistic to assume that the responses of exports to changes in the explanatory variables can be delayed as purchasers take time to readjust their spending patterns and also because of the delay between the placing of an order and delivery. The majority of the dynamic models use the Almon distributed lag pattern to specify the time shape of the reaction of the quantities exported to changes of relative prices, world activity and demand pressure variables (Aurikko (1975), Hutton and Minford (1975), Batchelor (1977a and 1977b), Laury and Warburton (1977), Richardson (1977)).

Most of the studies we refer to have been carried out at a level of aggregation dealing mainly with the exports of manufactures. Only Aurikko (1975) and Kwack (1972) work at a disaggregated level of five groups of commodities for the Swedish and U.S. exports respectively, whereas Winters (1976 and 1977) examines the export behaviour of sixteen commodity groups of U.K. exports.

Finally, as in the case of import demand equations, almost all the authors have estimated their models by ordinary least squares. An exception is Turnovsky (1968), who employs constrained least squares in order to incorporate non-linear constraints on the coefficients of competitors' export prices and world activity variable and their one-period lagged values.

CHAPTER IV

MODEL SELECTION PROCEDURES AND FORMULATION OF GREECE'S IMPORT AND EXPORT DEMAND EQUATIONS

As was mentioned in chapter II, Greece's import and export demand functions considered previously were static relationships in which all variables relate to a single time period; that is, it is assumed that consumers adjust to changed conditions within this single period. But in many cases, adjustment spreads over more than a single period and the demand function becomes a dynamic relationship, depending not only on the current level of its influences but on their past levels as well. Moreover, dynamic demand functions can be developed by including into the relationship the effects of past purchases of the good under consideration.

However, we should notice that the need to consider dynamic relationships does not depend only on the possible delay of the explanatory variables' effects, but also on the time aggregation of the data used. For example, if a six-month lag effect is suspected, it is difficult to detect it with an annual model. On the contrary, data disaggregated over time may cause dynamic effects, if the period between two successive observations is smaller than the time the effect of a variable takes to be demonstrated.

To test these hypotheses with observable data we must specify the time shape of the reaction of the quantity demanded, but on this economic theory has little to say and therefore a priori information is not available. Thus an empirical specification procedure has to be employed in order to face the problem of determining the dynamic specification of the model. The traditional selection procedure is to experiment with the techniques suggested by the distributed lag literature (e.g. Dhrymes (1971)) and this approach has been followed by those who have developed dynamic import and export demand functions for other countries, as mentioned in chapter III.

In the first section of this chapter we describe alternative methods for the empirical specification of dynamic models, their characteristics in relation with monthly data and their computational aspects. In section 2 we present the basic model we adopt for the Greek foreign trade which will serve as a baseline for our subsequent empirical analysis.

1. Model Selection Procedures

1.1. Starting from a Static Model

In cases where the traditional distributed lag techniques have been adopted for the empirical specification of dynamic models, tests of serial correlation in the residuals have been used to check the adequacy of the model and then serve as a guide to readjustments of the lag pattern of the explanatory variables.

However, such conventional procedures may be misleading for two main reasons. First, if we apply, in the presence of autocorrelated errors, the usual least-squares formulae for the sampling variances of the regression coefficients we are likely to obtain a serious underestimate of these variances, and so overestimated t ratios of the estimated coefficients. Second, if lagged values of the dependent variable appear among the regressors, the conventional Durbin - Watson test is biased towards the value for a random disturbance in these same circumstances (see Nerlove and Wallis (1966)). Thus the conventional criteria, on which the model is assessed, are likely to be unreliable. Although these difficulties can be avoided by using a test designed to cater for the lagged dependent variable case, e.g. Durbin's (1970) h test or a likelihood ratio test, and reestimating to take account of any autocorrelation detected, the above procedure is characterized by a lack of systematization in determining the dynamic specification.

Alternatively, starting from the static form of the relationship, serial correlation can be used as a convenient tool to specify the dynamic form of the relationship. Such an approach is based principally on the fact that misspecified dynamics may result in a serially correlated disturbance and considers tests which check this and allow us to discriminate between stochastic specifications and dynamic specifications. Moreover, as we shall see below, basing a procedure on the interaction between the equation dynamics and the stochastic specification allows us to extend the dynamic specification in a systematic way when the relevant tests indicate that the particular model estimated, at any stage of the procedure, is inadequate.

The latter approach has been considered by Sargan (1964) (see also Wallis (1972) and Hendry (1974 and 1978)), and it gains in popularity in applied econometric work due also to the increased availability of computing facilities which permit the estimation of most forms of models with either autoregressive or moving average errors (Hendry and Srba (1976 and 1977)).

To illustrate the above model selection procedure, consider the relationship.

(1) $$y_t = \sum_{i=1}^{k} a_i x_{it} + u_t \qquad (t = 1,\ldots,T).$$

For simplicity we allow u_t to follow a first order autoregressive process, i.e.

(2) $$u_t = \rho_1 u_{t-1} + \varepsilon_t$$

where the ε_t are independent random variables.

Combining equations (1) and (2) we get

(3) $$y_t = \rho_1 y_{t-1} + \sum_{i=1}^{k} a_i x_{it} - \sum_{i=1}^{k} \rho_1 a_i x_{i,t-1} + \varepsilon_t$$

which we denote the restricted transformed equation (RTE), since the 2k + 1 regression coefficients are functions of k + 1 parameters $(\rho_1, a_1,\ldots,a_k)$.

If the systematic part of (1) is misspecified and the omission of the lagged variables causes the autocorrelation, then the correct relationship has a dynamic form, regressing y_t on $y_{t-1}, x_{1t},\ldots,x_{kt}$ and $x_{1,t-1}, \ldots, x_{k,t-1}$. Thus the hypothesis to be estimated is

(4) $$y_t = \gamma_0 y_{t-1} + \sum_{i=1}^{k} \beta_i x_{it} + \sum_{i=1}^{k} \gamma_i x_{i,t-1} + e_t$$

where the e_t are independent random variables. (This assumes that no x - variables are redundant when lagged, i.e. there are no variables which when lagged are identical with, or linear combinations of other variables appearing in the equation, examples being the constant term, trend and seasonal dummy variables and x - variables which have their own lagged values appearing among the k explanatory variables in (1). If there are m such redundant variables labelled $x_{k-m+1}, \ldots, x_k$, their lagged values do not appear in the equation and (4) becomes

$$(5) \qquad y_t = \gamma_0 y_{t-1} + \sum_{i=1}^{k} \beta_i x_{it} + \sum_{i=1}^{k-m} \gamma_i x_{i,t-1} + e_t).$$

Since (4) (or (5)) is equivalent to ignoring the restrictions between the coefficients in (3) imposed by the autoregressive error specification, it is denoted the unrestricted transformed equation (UTE).

Equations (1) and (4) (or (5)) are estimable by ordinary least squares and (3) by non-linear least squares. If there is more than one endogenous variable in the equation, then (1) and (4) are estimated by instrumental variables and (3) by the autoregressive instrumental variables approach proposed by Sargan (1959). Denoting by S_1, S_3 and S_4 the residual sums of squares of the equations (1), (3) and (4) respectively, we can construct F or X^2 tests based on the likelihood ratio principle to discriminate between the three alternatives.

First we test the significance of $\hat{\rho}_1$ in (3) either by using $T \log_e(S_1/S_3) \sim X^2(1)$ or by the asymptotically equivalent t test on $\hat{\rho}_1$[1]. Then we test the autoregressive restrictions in (3) against (4) using $T \log_e(S_3/S_4) \sim X^2(n)$ where n is the number of restrictions imposed on (4) to obtain (3). Finally we test the significance of γ's in (4) either jointly by the standard F test on additional variables in (1) or individually by t tests.

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[1] In the least squares case we can use the Durbin - Watson statistic

If all the tests are nonsignificant we choose (1) with white noise errors. If the first test is significant and the second one insignificant then we choose the autoregressive error specification (3) (i.e. (1) + (2)). If all the tests are significant we choose (4) as a new form of the structural equation and we repeat again the same process assuming as before a first order autoregressive equation for the errors. Notice that if not all the γ's in (4) are significant, the new structural equation will contain the lagged values of those variables with the significant coefficients in the unrestricted equation. The process of modifying the form of the structural equation stops when the first or the second test or both are not significant.

However, there are cases where a higher order autoregressive scheme in the errors provides a relevant hypothesis. In particular, as Wallis (1972, p. 617) points out, when seasonally unadjusted data are being employed in order that one may attempt to explain seasonal variation in the dependent variable, along with other types of variation, by means of explanatory economic or seasonal variables, then the presence of non-systematic seasonal variation, or an incomplete accounting for seasonality by the regressors, will produce seasonal effects in the error term, with the possible consequence of higher order autocorrelation.

To consider these cases the above analysis can be generalised for higher order autoregressive errors

(6) $$u_t = \rho_\ell u_{t-\ell} + \varepsilon_t$$

or for the general ℓ^{th} order autoregressive process

(7) $$u_t = \sum_{i=1}^{\ell} \rho_i u_{t-i} + \varepsilon_t$$

The corresponding restricted transformed equations are obtained by applying the lag operators $(1 - \rho_\ell L^\ell)$ and $(1 - \rho_1 L - \rho_2 L^2 - \ldots - \rho_\ell L^\ell)$, respectively, to equation (1) and give

(8) $$y_t = \rho_\ell y_{t-\ell} + \sum_{i=1}^{k} a_i x_{it} - \sum_{i=1}^{k} \rho_\ell a_i x_{i,t-\ell} + \varepsilon_t$$

for the autoregressive specification (6) and

(9) $$y_t = \sum_{i=1}^{\ell} \rho_i y_{t-i} + \sum_{i=1}^{k} a_i x_{it} - \sum_{j=1}^{\ell} \sum_{i=1}^{k} \rho_j a_i x_{i,t-j} + \varepsilon_t$$

for the autoregressive specification (7) assuming that no redundant variables appear among the regressors.

Equation (8) ((9)) contains k+1 (k+ℓ) parameters which are estimated simultaneously by minimizing the residual sum of squares

$$\varphi = \sum_{t=\ell+1}^{T} e_t^2$$

with respect to the parameters in question. With the assumption that the ε_t are normally distributed random variables, the resulting estimates are asymptotically equivalent to maximum likelihood estimates (see Wallis (1972), p.p. 629-631).

Finally, equation (8) ((9)) has 2k+1 (ℓ+k+ℓ.k) coefficients which are subject to k (ℓ.k) non-linear restrictions which if ignored we obtain the corresponding unrestricted transformed equation. Here again, to discriminate between the three alternatives the above described likelihood ratio tests are applied.

1.2. Starting from a General Unrestricted Dynamic Model

The method described in the preceding section is a stepwise approach which allows one to increase successively the order of lags until the final dynamic form of the relationship is obtained. That is we start with the simplest model and we test if it is necessary to consider a more general one.

To begin with the most restricted hypothesis and then test sequentially for less restricted ones, is a procedure which is often used in practice, but with some unfavorable statistical properties. In particular, testing in the above procedure can be arbitrary in that the complete hypothesis testing framework is not specified since a maintained hypothesis need not be specified at the outset, simply being the first, of the sequence of hypotheses tested, less restricted one not rejected. Therefore it is difficult to say anything about the procedure's power and the analysis of its statistical power becomes complicated due to the fact that the test statistics of the above procedure are not independent (see Mizon (1977b, p. 1227)).

An alternative method is to begin with a general unrestricted model and then attempt to reduce the number of parameters needed to specify the data generation process. Now the maintained hypothesis is a general unrestricted dynamic model which includes the maximum number of lags, and tests are carried out to ascertain whether restricted versions of it are consistent with the data.

To illustrate this, consider a stable regression equation relating the dependent variable y_t to its own lagged values and to the current and lagged values of the regressors x.

(10) $$a(L)\ y_t = \sum_{i=1}^{k} \beta_i(L)\ x_{it} + \varepsilon_t ,$$

where $a(L)$, $\beta_1(L)$, ... ,$\beta_k(L)$ are scalar polynomials in the lag operator L of degree p, q_1, ... , q_k respectively and ε_t is a white noise error. The question now arising is whether the above k+1 polynomials which determine the overall dynamics in the equation, contain as a common factor a polynomial $\rho(L)$ of degree m, where $m \leq \min(p, q_1, \ldots, q_k)$. If that is the case, (10) becomes

$$\rho(L)\ a^*(L)\ y_t = \rho(L) \sum_{i=1}^{k} \beta_i^*(L)\ x_{it} + \varepsilon_t$$

or

$$a^*(L)\ y_t = \sum_{i=1}^{k} \beta_i^*(L)\ x_{it} + u_t$$

where

$$\rho(L) u_t = \varepsilon_t$$

The polynomials $a^*(L)$, $\beta_1^*(L)$, ..., $\beta_k^*(L)$ are of degree p-m, q_1-m, ..., q_k-m respectively and determine the "systematic dynamics", whereas $\rho(L)$ gives the "error dynamics".

The above analysis assumes that the degrees of the lag polynomials are known and in practice this will not be the case. Hendry and Mizon (1978 and 1978b) describe two approaches which could be adopted to solve this, both being two - stage decision procedures which commence from the most general unrestricted model in which the lengths of the longest lags must be specified a priori (e.g. p, q_1, ..., q_k in (10)).

In the first approach we conduct sequential tests as in

Anderson (1971, ch. 3.2), for reducing the order of the dynamics until a test value exceeding the chosen critical value is obtained. Then the equation so selected with the overall lag length at the value found during the first stage is examined for common factors. The second approach is to apply the common factor technique directly to the a priori specified unrestricted model and if m common roots are found, to simplify the model by testing whether any of the common roots are zero. This so called "common factor" analysis has been proposed by Sargan (1975, 1976, and 1977) (see also Wallis (1976)) who also describes the computer algorithm for implementing the common root tests by the use of Wald criteria. Two examples with respect to the dynamic specification of a demand for money function in U.K. and a demand function for consumer durables in Canada, are given by Hendry and Mizon (1978 and 1978b). Sargan (1977), also, applies the above analysis to specify the dynamic form of a wage equation in the U.K.

Briefly the approach of testing for common roots works as follows. Firstly we estimate the coefficients and their variance matrix in the general unrestricted dynamic model (10) by the desired method (ordinary least squares or instrumental variables). Denote the vector of the estimated coefficients by $\underset{\sim}{\hat{\delta}}$ and the estimated variance matrix by $\underset{\sim}{\hat{V}}$. Let now

$$\underset{\sim}{a}(L) = \left[a(L) \;\; \beta_1(L) \;\; \ldots \;\; \beta_k(L) \right]$$

and

$$\underset{\sim}{a}^*(L) = \left[a^*(L) \;\; \beta_1^*(L) \;\; \ldots \;\; \beta_k^*(L) \right]$$

be row vectors of the corresponding k+1 scalar polynomials.

The existence of a common factor $\rho(L)$ implies that

(11) $\quad \underset{\sim}{a}(L) = \rho(L)\ \underset{\sim}{a}^{*}(L)$

The restrictions in (11) impose a set of n non-linear restrictions on vector $\underset{\sim}{b}$ (n = m.k if no redundant variables appear among the regressors). If we represent the restrictions in vector form

$$\underset{\sim}{\Phi}(\underset{\sim}{b}) = \underset{\sim}{0}$$

where

$$\underset{\sim}{\Phi}(\underset{\sim}{b}) = \{\Phi_1(\underset{\sim}{b})\ \Phi_2(\underset{\sim}{b})\ \ldots\ \Phi_n(\underset{\sim}{b})\} \quad ,$$

the asymptotic variance matrix of $\underset{\sim}{\Phi}(\underset{\sim}{b})$ is given by

$$\underset{\sim}{S} = \left(\frac{\partial \underset{\sim}{\Phi}}{\partial \underset{\sim}{b}}\right) \underset{\sim}{V} \left(\frac{\partial \underset{\sim}{\Phi}}{\partial \underset{\sim}{b}}\right)' \ .$$

Therefore, in large samples and under the null hypothesis that the restrictions $\underset{\sim}{\Phi}(\underset{\sim}{b}) = \underset{\sim}{0}$ are valid, $\underset{\sim}{\Phi}(\hat{\underset{\sim}{b}})'\hat{\underset{\sim}{S}}^{-1}\underset{\sim}{\Phi}(\hat{\underset{\sim}{b}})$ will be distributed as a central chi - squared variate with n degrees of freedom. The relevant computer program, named COMFAC, has been developed by Sargan and Sylwestrowicz (1976) and is used with Hendry and Srba's programs GIVE and RALS.

Suppose we now consider some set of constraints $\underset{\sim}{\Phi}_1(\underset{\sim}{b}) = \underset{\sim}{0}$ forming the first n_1 constraints of the set of n constraints $\underset{\sim}{\Phi}(\underset{\sim}{b}) = \underset{\sim}{0}$. The corresponding Wald criterion is $W_1 = \underset{\sim}{\Phi}_1(\hat{\underset{\sim}{b}})'\hat{\underset{\sim}{S}}_1^{-1}\underset{\sim}{\Phi}_1(\hat{\underset{\sim}{b}})$, where $\underset{\sim}{S}_1 = (\partial\underset{\sim}{\Phi}_1/\partial\underset{\sim}{b})\ \underset{\sim}{V}\ (\partial\underset{\sim}{\Phi}_1/\partial\underset{\sim}{b})'$, which is asymptotically distributed as a X^2 with n_1 degrees of freedom. If W is the Wald criterion for the initial set of n constraints $\underset{\sim}{\Phi}(\underset{\sim}{b}) = \underset{\sim}{0}$, then $W - W_1$ is asymptotically distributed as X^2 with $n - n_1$ degrees of freedom independent of W_1.

So, if W_{m-1} and W_m are the Wald criteria which test the validity of the restrictions required for the existence of a common factor polynomial $\rho(L)$ of degree m-1 and m respectively, then if the first set of restrictions is valid, we can test the validity of the adittional restrictions required for one extra common root, using the criterion $W_m - W_{m-1}$. Hence, testing sequentially for common roots, we can either use the (overall) Wald criteria W_1, W_2, ... W_m or the difference Wald criteria W_1, W_2-W_1, W_3-W_2, ..., W_m-W_{m-1}. However, it is not easy to represent explicitly the set of constraints for the case where the degree of $\rho(L)$ is m, as made up of first the set of constraints for the case where the degree of $\rho(L)$ is m-1, plus a set of, say k, extra constraints (see also the example in section 1.4. below). Thus, as Sargan (1977) shows, if the true degree of the common factor polynomial is $\bar{m}$, then the above Wald test is asymptotically valid to test $m = \bar{m}$ against $m = \bar{m}+1$, but is not valid to test $m = \bar{m}$ against $m=\bar{m}-1$.

The contrast between the tests of specification of the above approach (which systematically tests restricted models within a general maintained hypothesis) and the tests of misspecification of the approach described in section 1.1. (which tests the need to consider more general models usually without a specified maintained hypothesis) has been emphasized by Mizon (1977b) who also discusses the statistical properties of the former approach (see also Mizon (1977a) and Hendry and Mizon (1978)). As mentioned before, the maintained hypothesis is the general unrestricted dynamic model

(12) $\quad \underset{\sim}{a}(L)\ \underset{\sim}{x}_t = \varepsilon_t$

where $\underset{\sim}{x}_t$ is the column vector $\{y_t\ x_{1t}\ \ldots\ x_{kt}\}$ and the restricted version of it has the form

(13) $\quad \rho(L)\ \underset{\sim}{a}^*(L)\ \underset{\sim}{x}_t = \varepsilon_t$

or equivalently

$$\underset{\sim}{a}^{*}(L)\ \underset{\sim}{x}_t = u_t$$

$$\rho(L)\ u_t = \varepsilon_t$$

To empirically determine the dynamic specification of the model it is necessary to determine the order of dynamics (i.e. the length of lag for each variable $\hat{p} \leq p$ and $\hat{q}_i \leq q_i$, $i = 1,\ldots, k$), and to test whether the factorization (11) is consistent with the data for some value of $\hat{m} \leq \min(\hat{p}, \hat{q}_1,\ldots, \hat{q}_k)$. This is equivalent to determining the orders of systematic dynamics p-m and q_i-m, i = 1,...,k, and the order of error dynamics m. Although all the hypotheses concerning m and p-m, q_i-m, i = 1,...,k, in (13) give models nested within (12) the testing of these hypotheses is complicated by their lack of unique ordering. However, as mentioned before, the overall problem consists of two sub-problems; first determine $\hat{p}$ and $\hat{q}_i$, $i = 1,\ldots,k$, and then determine $\hat{m} \leq \min(\hat{p}, \hat{q}_1,\ldots, \hat{q}_k)$ or alternatively first determine $\hat{m} \leq \min(p, q_1,\ldots, q_k)$ and then determine the systematic dynamics $p-\hat{m}$, $q_i-\hat{m}$, $i = 1,\ldots, k$. So we have two approaches each consisting of two subproblems and the hypotheses involved in each of these sub-problems are uniquely ordered.

If in (12) $p = q_1 = \ldots = q_k = q$ then the first stage of the above two-stage procedure consists of testing sequentially the following hypotheses:

$$H_i : \underset{\sim}{\beta}_{q-i} = \underset{\sim}{0} \quad \text{for } i = 0,1,\ldots, q$$

where $\underset{\sim}{\beta}_{q-i}$ is the vector of the coefficients of $\underset{\sim}{x}_{t-(q-i)}$ variables. The hypothesis accepted is the one immediately prior to that which produces a significant test criterion. These hypotheses are uniquely ordered and the above sequential procedure will have desirable properties analogous to those outlined by Anderson (1971). In particular, Anderson (1971) has shown that procedures which test sequentially hypotheses in increasing order of restrictiveness are uniformly most powerful in the class of procedures that fix the probabilities of accepting a less restricted hypothesis than the true one. We should notice that if the lag polynomials $a(L)$, $\beta_1(L)$, ... , $\beta_k(L)$ are not restricted to be of the same order then the above sequence of hypotheses can be applied for each lag polynomial separately. Although for each variable separately the sequence of tests will have the optimal properties discussed above, in general these tests for the different variables will not be independent nor is there an ordering among the corresponding hypotheses, and so a uniformly most powerful procedure is not possible for the whole of the first stage (see Mizon (1977a), p. 111).

If q^* is the maximum order of the lag polynomials determined in the first stage, then the second stage consists of testing the sequence of non-linear hypotheses

$$\rho_m(L)\ \underset{\sim}{a}^*_{q^*-m}(L) = \underset{\sim}{a}_{q^*}(L)$$

for $m = 0, 1, 2, \ldots, q^*$. Though the Anderson analysis was for normal linear models it is asymptotically valid for more general non-linear models. This asymptotic extension of the Anderson analysis relies essentially on the fact that the asymptotic distribution of the statistic for testing any hypothesis in the ordered sequence against the less restricted hypothesis immediately preceding it depends on the validity of all less restricted hypotheses in the sequence, but not on that of more restricted hypotheses, and that each of these test statistics is asymptotically independent (see also Sargan (1975)).

Considering now the alternative two-stage procedure we see that again each stage consists of a uniquely ordered sequence of hypotheses. In particular in the first stage we determine how many common factors are consistent with the data at the chosen significance level, within the maintained hypothesis. The ordered sequence of hypotheses is

$$\rho_m(L)\ \underset{\sim}{a}^{*}(L) = \underset{\sim}{a}(L)$$

for $m = 0, 1, 2, \ldots, \bar{q}$ where $\bar{q} = \min(p, q_1, \ldots, q_k)$. Then we test for zero roots among the set of m common roots extracted and the second stage consists of testing sequentially the following hypotheses:

$$H_1 : \rho_m = 0$$

$$H_2 : \rho_{m-1} = \rho_m = 0$$

$$\cdots\cdots\cdots\cdots\cdots\cdots$$

$$H_m : \rho_m = \rho_{m-1} = \cdots = \rho_1 = 0$$

where ρ_i is the i^{th} coefficient of the error polynomial $\rho(L)$.

So for each of the two stages within each procedure the hypotheses to be tested form a uniquely ordered sequence and the tests induced in each stage will have high power asymptotically. But, since the two stages are not independent the same cannot be said of the procedure as a whole, but for a wide range of models they will have good power properties even in small samples according to a simulation study carried out by Hendry and Mizon (1978a) where they examine the small sample properties of the Wald test of common factor restrictions.

In a sequential testing procedure as the above, an important consideration is the choice of Type I error probabilities for each test in the sequence. For example the first stage of the first procedure mentioned before, has a significance level $\alpha = 1 - (1-\varepsilon)^n$, where n is the maximum number of tests possible in the sequence (i.e. q) and ε is the significance level common to all tests in this sequence. In practice this point is often ignored and a series of tests is performed using conventional significance levels for each test, so that the overall significance level can be very large. However, to the extent that the consequences of inconsistency are believed to be more serious than those of inefficiency in estimation, the implicit choice of large significance levels might be reasonable.

Finally, it should be noted that when we test linear hypotheses, as for example in the first stage of the first described two-stage procedure where we test linear hypotheses about model (12) which is linear in parameters, testing using either the Wald or the likelihood ratio principle will yield identical results. This is not the case when we test non-linear hypotheses, as for example in the second stage of the above procedure, but statistically the two principles will lead to asymptotically equivalent tests. However the Wald testing principle is computationally less expensive because estimation of the parameters corresponding to all hypotheses, as required by likelihood - ratio tests, is not necessary.

1.3. Seasonality

When quarterly or monthly data are employed for the estimation of economic relationships, the problem of seasonal variation is introduced. In this section we present the various methods which have been proposed to deal with the problem of seasonality and we discuss the problems which may arise from their application.

Generally, there are two ways to deal with seasonality. Either to work with data which have already been seasonally adjusted, or to use seasonally unadjusted data and to estimate the seasonal variation within the model, where the most common procedure is to add dummy (zero-one) variables to the model. When the first approach is adopted there are two alternatives. The first one is, assuming an additive components model, to remove the seasonal variation from the data, using a regression model of seasonal variation. If the seasonal pattern is constant over time, then the seasonal adjustment procedure can be carried out by regressing the unadjusted series on zero-one dummy variables, equal in number to the number of seasons per year. In this case the resulting coefficients are the seasonal means and the adjusted series x_t^a is obtained by adding the overall mean of the series x_t to the regression residuals. The second alternative is to use non-linear adjustment procedures assuming that the trend-cycle, seasonal and irregular components are connected with a multiplicative model. Such a procedure is the "ratio-to-moving-average" technique on which the official adjustment procedures are largely based.

Suppose that $\underset{\sim}{y}$ is the vector of observations on the dependent variable, $\underset{\sim}{X}$ is the matrix of explanatory variables and $\underset{\sim}{D}$ is the matrix of the seasonal zero-one dummies. Presume also, that $\underset{\sim}{y}^a$ and $\underset{\sim}{X}^a$ are the corresponding seasonally adjusted series obtained by the above described "least-squares adjustment" technique.

Lovell (1963) has shown that the least squares estimates of the vector of x coefficients are the same when:

(i) $\underset{\sim}{y}$ is regressed on $\underset{\sim}{X}$ and $\underset{\sim}{D}$

(ii) $\underset{\sim}{y}$ is regressed on $\underset{\sim}{X}^a$

(iii) $\underset{\sim}{y}$ is regressed on $\underset{\sim}{X}^a$ and $\underset{\sim}{D}$

(iv) $\underset{\sim}{y}^a$ is regressed on $\underset{\sim}{X}^a$

(v) $\underset{\sim}{y}^a$ is regressed on $\underset{\sim}{X}$ and $\underset{\sim}{D}$

However, Thomas and Wallis (1971) suggest that model (i) is preferred because, in contrast with the others, it takes explicitly into consideration the loss of degrees of freedom due to the estimation of the D's coefficients. They, also, show that if the model is misspecified so that the seasonal variables are erroneously introduced, then the least squares estimates of the X's coefficients are inefficient. On the contrary if misspecification is considered in the opposite case so that the seasonal variables are erroneously excluded from the model, then the estimated coefficients are biased. Both results are not valid if the X-variables are uncorrelated with the seasonal dummy variables. So, to avoid any bias in the estimation of the structural coefficients caused by the erroneous exclusion of the seasonal variables, it is preferable to choose one of the equivalent (i) - (v) forms with the risk of obtaining inefficient estimates but unbiased ones. Of course, a test of the hypothesis that the D's coefficients are zero is possible and below we suggest the stage of the dynamic specification procedure, at which this test should be carried out and also which of the (i) - (v) forms should be adopted in our subsequent empirical analysis.

The effects of seasonal adjustment procedures, based on the ratio-to-moving-average method, on the relations between variables have been examined by Wallis (1974). He demonstrates that the widely used U.S. Bureau of the Census Method II, Variant X-11 procedure can be approximated by a two-sided lag polynomial of the form $A(L) = \sum_{j=-m}^{m} a_j L^j$ ($a_j = a_{-j}$) and hence seasonally adjusted data take the form $y_t^a = A(L) y_t$.

If two series y_t and x_t are related by the distributed lag model

$$y_t = \beta(L)\, x_t + u_t$$

and are seasonally adjusted using linear filters (as $A(L)$ can be interpreted) $A_y(L)$ and $A_x(L)$ such that $y_t^a = A_y(L) y_t$ and $x_t^a = A_x(L) x_t$, then the relation between the adjusted variables is

$$y_t^a = \frac{A_y(L)\ \beta(L)}{A_x(L)}\ x_t^a + A_y(L)\ u_t$$

It is seen that the lag function is distorted and the error term is changed to $u_t^a = A_y(L)\ u_t$. If the same linear filter adjustment procedure is applied to both series (i.e. $A_y(L) =$

$A_x(L)$), then the relationship between them is not changed and the only effect is on the error term. In this case the transformed error $A_y(L)\ u_t$ will be white noise only if u_t is seasonal and the filter used for y_t eliminates the original seasonal noise in u_t. Then as a result the adjusted data will provide more efficient estimates. Otherwise, as Wallis (1974) shows, the above seasonal adjustment procedure can create serial correlation or distort the existing pattern, and therefore dynamic specification problems may arise in the context of the dynamic specification approach described in section 1.1.

Our data are monthly, seasonally unadjusted, series from 1954 - 1976 and in view of the above discussion it was decided that one of the equivalent (i) - (v) forms, should be used to take account of the seasonal variation present in the monthly series. Forms (i), (iii) and (v) imply that eleven seasonal shift variables plus a constant term are to be included in every equation. But, one of the two programs of Hendry we use, required enormous dimensions (see Hendry and Srba (1977) for details) when we allowed the error term to follow a 13th order general autoregressive process, as we shall see below[1].

---

[1] As an indication we mention that one of the arrays the program requires is of order $T.\{(k+1).(\rho+1)\}$, where T is the sample size, k is the number of explanatory variables and ρ is the order of the autoregressive process. For $T=276$, $k=14$ (including eleven seasonal dummies, one constant term and two economic explanatory variables) and $\rho=13$ the above array requires $(276\times210) = 57{,}960$ elements (58 K).

So one of the two alternative approaches (ii) and (iv) should be adopted which reduce the number of regressors by eleven. Lovell (1963, p. 1001) has shown that equation (iv) gives the same regression residuals with equation (i), whereas the regression residuals of equation (ii) are equal to the residuals of regression (i) plus the difference between the unadjusted and adjusted series y_t (i.e. $e_{(ii)} = e_{(i)} + (y - y^a)$). That leads us to choose form (iv) which gives the same residual sum of squares as equation (i) and allows us at any stage of the dynamic specification procedure to test if the seasonal adjustment of the series is necessary. Moreover, the dynamic specification approach based on the likelihood ratio principle remains unaffected.

Therefore, a computationally inexpensive approach is the following. We seasonally adjust the series using the "least - squares adjustment" technique and then select the preferred specification according to the likelihood - ratio or the Wald criteria principle. Then we reestimate the same equation using unadjusted data and from the corresponding residual sums of squares we construct the F - test described by Thomas and Wallis (1971, p. 63) to test the hypothesis that in the equivalent (but not estimated) form (i) the coefficients of D's are zero. Rejection of the null hypothesis implies that seasonal variables are required to explain the seasonal variation in $\underset{\sim}{y}$ and the equivalent least - squares adjustment of the series was necessary. But if the null hypothesis is accepted then the adjustment of the series is redundant and the preferred form is the equation estimated from the unadjusted series. However, in order to make sure that the preferred form is the correct one, we should repeat the dynamic specification procedure using this time unadjusted data, since, in general, the regression residuals obtained from the unadjusted series do not exhibit the same behaviour as those obtained from the adjusted series.

1.4. Application of the Model Selection Procedures to Monthly Data

The use of monthly data from the period 1954 - 1976 instead of annual or quarterly observations, which have been employed in previous work on models of foreign trade, constitutes a novelty of our approach. In the context of monthly data we describe in this section the application of the empirical dynamic specification procedures discussed above.

Let us consider first the approach which starts with the simplest (static) model. As already mentioned, tests for the presence of autocorrelation serve as tests of misspecification in order to check whether lagged values of the variables should be included into the equation. In economic models, the error term is usually taken to represent the effects of omitted, unobservable variables on the dependent variable, and in the context of seasonal variation in the monthly series, these omitted variables may themselves possess seasonal variation in one form or another. In particular, to the extent that the seasonal variation within the model is not exactly systematic, and may not be completely accounted for by whatever seasonal variables have been included, then these effects will appear in the error term. So this kind of seasonal variation may produce autocorrelation in the error term and with monthly data one would expect to find a relationship between the residuals in the corresponding months of successive years. This is one reason to test for simple autoregressive models of the form

(1) $$u_t = \rho_{12}\, u_{t-12} + \varepsilon_t$$

On the other hand if a dynamic form of the relationship is more consistent with the data, then with monthly series and with the presence of seasonal factors, the omission of the twelve month lagged variables should be considered as a possible source of misspecification. Of course, this can be tested with the likelihood ratio test described before, if the autoregressive coefficient in (1) is found to be significant. Therefore, we see that a 12-month lag effect either in the error term or in the systematic part of the structural equation has to be considered when one is working with monthly figures. However, nothing is known a priori about the form of the autoregressive process in the errors, or equivalently, about the correct dynamic form of the equation, so that the omission of the lagged variables causes the autocorrelation. All this is undertaken in an exploratory spirit and so in our subsequent empirical analysis we consider all the alternative forms of the model

(2) $$u_t = \rho_\ell\, u_{t-\ell} + \varepsilon_t$$

for $\ell = 1, 2, \ldots, 12$.

The above procedure allows us, using the corresponding likelihood ratio tests, to discriminate between the original equation, simple autoregressive error specification and the corresponding unrestricted dynamic form, if ρ_ℓ in (2) is found significant. But it is likely to obtain significant results for more than one values of ℓ and that indicates that the error term may follow a more general autoregressive process or, alternatively, that different lags of the dependent and the explanatory variables should be included into the equation to obtain the proper dynamic form of the relationship. To test these hypotheses we have also allowed the error term to follow the general autoregressive scheme

$$(3) \qquad u_t = \sum_{i=1}^{\ell} \rho_i u_{t-i} + \varepsilon_t$$

for $\ell = 1, 2, \ldots, 12$.

The maximum likelihood estimation of (2) and (3) permits various likelihood ratio tests on the autoregressive parameters, jointly or singly, to be carried out. For example, to choose between (2) and (3) for some value of ℓ, say 12, we use the likelihood ratio test $T \log (S_2/S_3) \sim X^2(11)$, where S_2 and S_3 are the corresponding residual sums of squares, to test the hypothesis $H_0 : \rho_1 = \rho_2 = \ldots = \rho_{11} = 0$. Then, if S_R is the residual sum of squares of the preferred restricted form and S_U is the residual sum of squares of the corresponding unrestricted (dynamic) form, to discriminate between them we use the likelihood ratio test $T \log (S_R/S_U) \sim X^2(n)$ which tests the validity of the n restrictions imposed on the coefficients of the unrestricted equation to obtain the restricted one.

An alternative autoregressive error specification which we also consider, is the multiplication of the first order and twelfth order quasi-difference operators (see Wallis (1972),

p. 629), namely,

(4) $$(1 - \rho_1 L)\ (1 - \rho_{12} L^{12})\ u_t = \varepsilon_t$$

where L is the lag operator. Specification (4) can be interpreted either as a simple 12th order autoregressive process of the transformed error $v_t = u_t - \rho_1 u_{t-1}$, or as a first order autoregressive scheme of the transformed error w_t, where $w_t = u_t - \rho_{12} u_{t-12}$. The autoregressive error structure (4) is preferred if the hypothesis that $\rho_{13} = -\rho_1\ \rho_{12}$ in the model

(5) $$u_t = \rho_1\ u_{t-1} + \rho_{12}\ u_{t-12} + \rho_{13}\ u_{t-13} + \varepsilon_t$$

is accepted, or more generally, if in the model (3) with $\ell=13$ the joint hypothesis $\rho_2 = \rho_3 = \ldots = \rho_{11} = 0$, $\rho_{13} = -\rho_1 \rho_{12}$ is accepted. Here again the validity of the restrictions is tested by the above described likelihood ratio test, where this time S_R is the residual sum of squares of specification (4) (or (5)) and S_U is the residual sum of squares of the unrestricted structural equation which contains one, twelve and thirteen month lagged values of all variables.

The above empirical specification approach which includes the estimation of all the already mentioned restricted and unrestricted equations, has been applied for every import and export demand equation we consider. But, to save space, only the forms which produced significant results are presented, as we shall see, below.

Considering now the "common factor" analysis, we require as a baseline a general unrestricted dynamic model of the form

(6) $$a(L)\ y_t = \sum_{i=1}^{k} \beta_i(L)\ x_{it} + \varepsilon_t$$

where $a(L)$, $\beta_1(L)$, ..., $\beta_k(L)$ are scalar polynomials in the lag operator L and ε_t is a white noise error. As mentioned in section 1.2. two approaches can be adopted to determine the orders of the lag polynomials in (6). Since our data are monthly series and in view of the above discussion, we have adopted the second approach (which applies directly the common factor technique to the a priori specified unrestricted model) in our empirical analysis, and we allowed the scalar polynomials in (6) to be of order 13. Then the common factor technique is applied and the equation is simplified by testing for zero roots from the set of r common roots extracted.

The ε_t in (6) are assumed to be independent random variables and to test this the value of the Box - Pierce test statistic for a random residual correlogram is also computed (see Pierce (1971)).

As mentioned before, equation (6) is easily estimated but the restrictions on the coefficients and their derivatives can be complicated to compute according to the degree of the common factor polynomial we test for. For example, consider the following unrestricted dynamic model

(7) $$y_t = a_1 y_{t-1} + a_2 y_{t-2} + \beta_0 x_t + \beta_1 x_{t-1} + \beta_2 x_{t-2} + \varepsilon_t$$

If a second degree common factor polynomial exists, so that the true model is

$$y_t = \beta_0 \, x_t + u_t$$

$$u_t = \rho_1 \, u_{t-1} + \rho_2 \, u_{t-2} + \varepsilon_t$$

the implied restrictions are

$$a_1 \, \beta_0 + \beta_1 = 0 \quad \text{and} \quad a_2 \, \beta_0 + \beta_2 = 0$$

Suppose now that we test for one common root so that the model

$$y_t = \gamma_0 y_{t-1} + \gamma_1 x_t + \gamma_2 x_{t-1} + u_t$$

$$u_t = \rho u_{t-1} + \varepsilon_t$$

is the data generation process (see Hendry and Mizon (1978)). The restricted transformed equation is

$$y_t = (\rho+\gamma_0)y_{t-1} - \rho\gamma_0 y_{t-2} + \gamma_1 x_t + (\gamma_2-\rho\gamma_1)x_{t-1} - \\ - \rho\gamma_2 x_{t-2} + \varepsilon_t$$

which has four parameters to be estimated as compared with the unrestricted model (7) which has five. Hence there is only one restriction to be tested, but it is more difficult to parameterise, in comparison with the previous restrictions, since the mapping between the restricted and unrestricted parameters

$$a_1 = \rho+\gamma_0 \qquad \beta_0 = \gamma_1 \qquad \beta_2 = -\rho\gamma_2$$

$$a_2 = -\rho\gamma_0 \qquad \beta_1 = \gamma_2 - \rho\gamma_1$$

implies the restriction

$$\frac{a_2^2\beta_0^2\beta_2+\beta_2^3+2a_2\beta_0\beta_2^2+a_1\beta_1\beta_2^2-a_2\beta_1^2\beta_2+a_1^2\beta_0\beta_2^2-a_1a_2\beta_0\beta_1\beta_2}{(a_2\beta_0 + \beta_2)^2} = 0$$

from which dividing by $-\beta_2$, assuming that β_2 is non-zero, the simplified numerator gives

$$-a_2^2\beta_0^2-\beta_2^2-2a_2\beta_0\beta_2-a_1\beta_1\beta_2+a_2\beta_1^2-a_1^2\beta_0\beta_2+a_1a_2\beta_0\beta_1 = 0$$

Therefore even for this second order dynamics example the calculation of the Wald test statistic for one common factor is complicated, and the degree of complexity increases with the order of dynamics and the number of common factors being tested. These considerations led Sargan (1975, and 1977) to note that the constraints on the coefficients of equation

(6) can be expressed in the form that a certain matrix, whose elements consist of the coefficients of (6) and zeros, should have a given rank, which can in turn be translated into equivalent determinental conditions.

To illustrate this, consider again the unrestricted dynamic model (7). Let $\underset{\sim}{A}_0$ be a matrix of order 2×3 whose elements are the coefficients of the two lag polynomials, i.e.

$$\underset{\sim}{A}_0 = \begin{bmatrix} 1 & -a_1 & -a_2 \\ \beta_0 & \beta_1 & \beta_2 \end{bmatrix}$$

According to Sargan, a necessary and sufficient condition that the two polynomials have a common factor of degree two, is that $\underset{\sim}{A}_0$ is of rank one which leads to the same restrictions as before, namely

$$\beta_1 + \beta_0 a_1 = 0 \quad \text{and} \quad \beta_2 + \beta_0 a_2 = 0$$

Consider now the matrix

$$\underset{\sim}{A}_1 = \begin{bmatrix} \underset{\sim}{A}_0 & \underset{\sim}{0}_{21} \\ \underset{\sim}{0}_{21} & \underset{\sim}{A}_0 \end{bmatrix}$$

where $\underset{\sim}{0}_{21}$ represents a zero matrix of dimensions 2×1. A necessary and sufficient condition that the two polynomials in (7) have a common factor of degree one is that $\underset{\sim}{A}_1$ is of rank three. This implies that $|\underset{\sim}{A}_1| = 0$ and gives

$$-a_2^2\beta_0^2-\beta_2^2-2a_2\beta_0\beta_2-a_1\beta_1\beta_2+a_2\beta_1^2-a_1^2\beta_0\beta_2+a_1a_2\beta_0\beta_1 = 0$$

which is the same as the restriction previously derived.

Another point to be mentioned is that the indication from the Wald test that a common factor polynomial exists suggests nothing about the autoregressive form of the error term. If, for example, the Wald test shows that there is a common factor polynomial of degree ℓ, and assuming that all the ℓ common roots are non-zero, then the error term may follow a simple autoregressive model of the form (2) or a general one of the form (3). The real form of the autoregressive process will depend on the values of the common roots. So in cases where the Wald test indicates that the lag polynomials in (6) contain a common factor of degree 13, and with the assumption that all the common roots are non-zero, the common factor polynomial can be of the form (2), (3), for $\ell=13$, (4) or (5). The exact form of the autoregressive process will be determined after the maximum likelihood estimates of these forms have been obtained and the already mentioned likelihood ratio tests have been carried out.

Consider now the case where the lag polynomials $a(L)$, $\beta_1(L)$, ..., $\beta_k(L)$ in (6) are of degree ℓ and the Wald test shows that there is a common factor polynomial of degree ℓ. Assume also that the true model is of a static form with an error term following a simple autoregressive process of order ℓ which has been confirmed by the likelihood ratio tests. In such a case one should expect the coefficients of the variables lagged one up to $\ell-1$ periods to be non-significant and in a F or X^2 test the unrestricted transformed equation which contains only the current and ℓ periods lagged values of all the variables to be accepted against the general unrestricted equation (6). But in practice this will not be always the case as our empirical analysis demonstrates. However these tests may give some indication about the possible form of the autoregressive process.

Considering now the error specification (4) (or (5)) we see that the corresponding unrestricted transformed equation contains the current and one, twelve and thirteen months lagged values of all the variables. Hence, if the common factor analysis indicates that equation (6), for $\ell=13$, contains a common factor polynomial of degree 13 then we can distinguish the two unrestricted forms using a F or a X^2 test. If the former unrestricted model is accepted we can test the validity of the non-linear restrictions imposed on its coefficients by the error specification (4) (or (5)). The test is based on the Wald testing principle and is computed by an algorithm called REST developed by Sargan and Sylwestrowicz (1976), for use with the programs RALS and GIVE. This subroutine computes the approximate Wald criteria for any set of constraints provided that they are programmed by the user (see Sargan and Sylwestrowicz (1976) for details).

For one regressor the above unrestricted transformed equation is

$$y_t = a_1 y_{t-1} + a_{12} y_{t-12} + a_{13} y_{t-13} + \beta_0 x_t + \beta_1 x_{t-1} + \beta_{12} x_{t-12} + \beta_{13} x_{t-13} + \varepsilon_t$$

or

$$(1-a_1 L - a_{12} L^{12} - a_{13} L^{13}) y_t = \beta_0 (1 + \frac{\beta_1}{\beta_0} L + \frac{\beta_{12}}{\beta_0} L^{12} + \frac{\beta_{13}}{\beta_0} L^{13}) x_t + \varepsilon_t$$

Specification (5) implies the restrictions

$$a_1 \beta_0 + \beta_1 = 0 \ , \quad a_{12} \beta_0 + \beta_{12} = 0 \quad \text{and} \quad a_{13} \beta_0 + \beta_{13} = 0 \ ,$$

whereas specification (4) requires one extra restriction for the factorization of the two lag polynomials, namely

$$(1-a_1L)(1-a_{12}L^{12})y_t=\beta_0(1+\frac{\beta_1}{\beta_0}L)(1+\frac{\beta_{12}}{\beta_0}L^{12})x_t + \varepsilon_t$$

which in view of the above restrictions implies that

$$a_{13}+ a_1a_{12} = 0 \quad \text{or} \quad \beta_{13} - \frac{\beta_1\beta_{12}}{\beta_0} = 0 \; .$$

Alternatively, specification (4) requires first the factorization of the two lag polynomials and in effect the restrictions

$$a_{13} + a_1a_{12} = 0 \quad \text{and} \quad \beta_{13} - \frac{\beta_1\beta_{12}}{\beta_0} = 0 \; ,$$

then, if valid,we might test for common factors implying the restrictions

$$a_1\beta_0 + \beta_1 = 0 \quad \text{and} \quad a_{12}\beta_0 + \beta_{12} = 0 \; .$$

So, even though the above restrictions are easily parameterized they are not uniquely ordered and the suggested procedure will not be powerful. Moreover, as the number of regressors increases, the number of alternative orderings increases too. However, since the COMFAC and REST algorithms are computationally inexpensive and easy to use the above Wald tests can provide useful information about the possible form of the error dynamics.

1.5. Computational Aspects

The above described empirical specification procedures require, generally, the estimation of the model

$$(1) \qquad y_t = \sum_{i=1}^{k} a_i \, x_{it} + \sum_{j=1}^{m} a_{k+j} \, y_{t-j} + u_t \quad (t = 1,\ldots,T)$$

where

(2) $$u_t = \rho_\ell \, u_{t-\ell} + \varepsilon_t$$

or

(3) $$u_t = \sum_{i=1}^{\ell} \rho_i \, u_{t-i} + \varepsilon_t$$

and the ε_t are independent random variables. When $\ell=0$, equation (1) is estimated by ordinary least squares, whereas for $\ell>0$ equations (1) and (2) (or (3)) are estimable by non-linear least squares. If there is more than one endogenous variable in the equation, then for $\ell=0$ equation (1) is estimated by instrumental variables and for $\ell>0$ (1) and (2) (or (3)) are estimable by the autoregressive generalization of the instrumental variable approach. Two computer programs have been used for the above estimation procedures, both developed by Hendry (1970) and described by Hendry and Srba (1976 and 1977). The first one, named GIVE (General Instrumental Variables Estimation of Linear Equations with Lagged Dependent Variables and First Order Autoregressive Errors), deals with cases where the error term follows a simple autoregressive process (i.e. (2)). The second computer program is called RALS (Rth-Order Autoregressive Least Squares) and estimates models with autoregressive errors of the form (3)[1].

The main problem with such iterative estimation methods is the possibility of the existence of several local minima in which case the process would converge arbitrarily to one of them, depending on the starting point. Program GIVE deals with this problem by calculating first the residual sum of

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[1] GIVE and RALS are programs from the AUTOREG Library: see Hendry and Srba (1978).

squares for a grid of values of ρ_ℓ from -0.92 to 0.98 in steps of 0.1 . This provides a rough check for multiple minima and helps ensure that the iteration commences close to the global minimum. A variant of Gauss-Seidel is then used till convergence is achieved. But in RALS this precaution against several local minima is not possible due to the biger number of autoregressive parameters to be estimated. The estimation process stops when the convergence criteria are met and a minimum is achieved without knowing whether it is a local minimum or a global one. With a few of our equations we faced this problem when we used logarithms of the variables and the error term followed either a general autoregressive scheme of order higher than three, or the factorised error specification

(4) $$(1 - \rho_1 L)(1 - \rho_{12} L^{12})\ u_t = \varepsilon_t$$

which RALS also estimates . For these cases, where all the minimization algorithms which are utilised by RALS, failed to produce acceptable results, the linear form was adopted. But in some cases it was impossible to estimate (4) even if both linear and log-linear forms were tried out, and so we were not able to consider this error specification in the empirical analysis of these particular equations.

GIVE and RALS are written in FORTRAN IV and were developed to operate on the University of London CDC 7600. After some alterations, we used them on the University of Warwick Burroughs 6700 whose limited capacity faced us with another problem created by the enormous process time required for every complete execution. In particular, GIVE proved to be quite fast and the required process time for the estimation of one equation including the original, restricted and unrestricted equations, was in all circumstances less than 30 seconds. But, RALS estimates every equation in a stepwise manner. If an ℓth order general autoregressive scheme is specified for a particular equation, the program first calculates the zero case, then the first order case, then the second order, etc. up to and including the ℓth order case itself. This estimation procedure required more than 3600 seconds (maximum permitted process time) when we attempted to use seasonal (zero-one) dummy variables and a thirteen order autoregressive scheme was specified, and it took 4000 seconds to reach only the 9th order case. To overcome this difficulty we tried to change the program so that an ℓth order autoregressive process was estimated straightaway, skipping the first up to ℓ-1 order cases (usually we were interested in estimating (3) for ℓ=12 and 13 but not for ℓ=1, ..., 11). But again the required process time exceeded the above limit. The execution time varies with the number of the regressors included in the equation and in view of the above difficulty it was decided to use deseasonalised data, apart from the problem of the size of the required dimensions, as mentioned in section 1.3.

The table below shows the time required for the estimation of each of the above error specifications, in a case where the structural equation included four regressors (two economic explanatory variables, one dummy and a constant term), and the sample consisted of 263 observations.

TABLE 1

PROCESS TIME REQUIRED ON THE BURROUGHS 6700

Error Specification	Program Used	Process Time (seconds)
$u_t = \rho_{12} u_{t-12} + \varepsilon_t$	GIVE	17
$u_t = \sum_{i=1}^{\ell} \rho_i u_{t-i} + \varepsilon_t$	RALS	
(ℓ = 12 and 13)		
- NAG Library's Routine E04AAF		476
- Powell's Algorithm		424
$(1-\rho_1 L)(1-\rho_{12}L^{12})u_t=\varepsilon_t$ and	RALS	482
$u_t=\rho_1 u_{t-1}+\rho_{12}u_{t-12}+\rho_{13}u_{t-13}+\varepsilon_t$		
(Powell's Algorithm)[1]		

The two different algorithms (see Hendry and Srba (1977) for details) used for the estimation of (3) required almost the same time and the resulting estimates were quite similar, but the NAG Library's algorithm proved to achieve a slightly lower residual sum of squares and so it was used for all our estimations. Specification (4) is non-linear in the auto-regressive parameters and its estimation involves the calculation of numerical second derivatives and therefore the required process time is significantly increased.

Another point which should be stressed is the possible bias in the computation of the likelihood ratio test, described before in section 1.4., to discriminate between the error

[1] Error specification (4) is estimated only by Powell's algorithm.

specifications (2) and (3) (or between (2) and (4)). As mentioned before, form (2) is estimated by GIVE and forms (3) and (4) are estimated by RALS. Since the minimization algorithms utilized by the two programs are different, the resulting residual sums of squares are also different. In particular, experiments made with first order autoregressive schemes, which is the only error specification estimable by both programs, showed that GIVE's algorithm achieves a lower residual sum of squares. Therefore the computed value of the likelihood ratio test, which refers to the hypothesis that in model (3) $\rho_1 = \rho_2 = \ldots = \rho_{\ell-1} = 0$ so that form (2) is preferred against (3), is biased downwards and should be accepted with a certain amount of caution, especially when it is near the critical value (the same is valid when we test the hypothesis that in (4) $\rho_1=0$ so that (2), for $\ell=12$, is preferred against (4)).

A final point to be mentioned is the number of observations included in our sample. The sample runs from January 1954 to December 1976 giving 276 observations on each variable. The different equations we have estimated contain up to thirteen months lagged values of the variables and therefore the series begin with the fourteenth observation reducing the sample size to 263 observations. In order to be able to carry out likelihood ratio tests to discriminate between the different forms, we have estimated all the equations using the last 263 observations of our sample.

2. Formulation of Greece's Import and Export Demand Equations

2.1. The Formulated Hypothesis

According to the preceding analysis we require a basic model which will serve as a baseline for the empirical dynamic specification procedure.

In perfectly competitive models, imports equal domestic demand less domestic supply. This implies that the imported and home produced commodities are identical (or homogeneous). Then the demand for imports is treated as the excess demand function of traditional international trade theory. Imports are regarded as the difference between two variables, total consumption and domestic production. Functions explaining these variables are formulated and the import function is the difference between these demand and supply functions. In other words, imports are regarded as not being demanded for their own sake, but because domestic supplies are not sufficient to meet demands at the prevailing price. No independent income elasticity is estimated for the imports and their price elasticities are derived from those of domestic supply and demand according to the formula

$$e_m = \frac{D}{M} e_d + \frac{S}{M} e_s$$

where D refers to domestic consumption, S to domestic supply and M to imports, while e_d and e_s are the price elasticities of demand and supply and e_m the import demand elasticity.

This formula is derived as follows: since the import-demand is the difference between domestic demand and domestic supply at the price P, the equation for import demand elasticity is

$$e_m = \frac{d(D-S)}{dP} \cdot \frac{-P}{D-S} = \frac{-P}{D-S} \cdot \frac{dD}{dP} + \frac{P}{D-S} \cdot \frac{dS}{dP}$$

Now multiply and divide the first component by D and the second component by S:

$$e_m = \frac{D}{D-S} \cdot \frac{dD}{dP} \cdot \frac{-P}{D} + \frac{S}{D-S} \cdot \frac{dS}{dP} \cdot \frac{P}{S} =$$

$$= \frac{D}{D-S} e_d + \frac{S}{D-S} e_s = \frac{D}{M} e_d + \frac{S}{M} e_s$$

(see Kreinin (1967), p. 515, and Scott (1963), p.p. 88-89 for a similar derivation).

But, if the imported commodity is not produced at home or the imported and home produced commodities are not identical, which is the usual case for Greece, the import demand coincides with the home demand for that commodity (clearly from the above relation if $S = 0$, $e_m = e_d$). Therefore, the problem of estimating the Greek import demand equation for a commodity or a group of commodities is generally reduced to the problem of estimating the country's home demand equation for such commodities.

The theoretical hypothesis, used in estimating import demand functions for commodities that are objects of final consumption, is provided by the theory of consumer's behavior, according to which the individual consumer allocates his income among consumable commodities in an effort to achieve maximum satisfaction.

The theory of consumer's demand suggests that the demand for a commodity by an individual consumer is a function of all commodity prices and his income. Extending that to the case of imports, we see that the quantity of imports purchased by any consumer will depend on his income, the price of imports, and the price of other consumable commodities, which in this case are domestic commodities.

Microeconomic theory of consumer's behavior deduces the demand functions of the individual consumer. In order to estimate them, we must observe the market behavior of the individual consumer. But such observations are not available and therefore estimation of individual demand functions is not feasible. Data are provided on an aggregative basis, and so we must formulate aggregate demand functions obtained by aggregating the demand relations for individual consumers over individuals and over commodities. This suggests that for an economy the import demand function for a group of commodities can be written as

$$M = f\,(P_m, P_d, I_G)$$

where

M = Quantity of the imported group of commodities

P_m = Price index of imported group of commodities

P_d = Index of domestic prices, and

I_G = Real income or activity variable (of Greece)

In a similar way we are led to the import demand function for a group of producer commodities, following the theory of the firm, according to which an enterpreneur attempts to produce that output for which his profits will be a maximum.

The theory of demand proceeds one step further in suggesting that the above demand relationship may actually be written as

$$M = f\left(\frac{P_m}{P_d}, I_G\right) .$$

This transformation is based on the assumption that individual consumers display no money illusion; that is, a doubling of all prices and money income will leave the quantity demanded unchanged. Moreover, this form has the statistical advantage that a possible collinearity between import and domestic prices, is avoided. It is also expected that

$$\partial M/\partial\left(\frac{P_m}{P_d}\right) < 0 \quad \text{and} \quad \partial M/\partial I_G > 0$$

So, economic theory suggests that the import demand for a group of commodities is a function of relative prices and income, and that a number of restrictions is imposed upon the signs of the relevant parameters. In order to fit such a relationship statistically using a relevant estimation method, a particular functional form must be chosen. But the theory has little to say about the exact mathematical form of the demand function. The most common forms are linear, as in equation (1)

$$(1) \qquad M = a_0 + a_1 I_G + a_2 \frac{P_m}{P_d} + u_t$$

and log-linear, as in equation (2)

$$(2) \qquad \log M = \beta_0 + \beta_1 \log I_G + \beta_2 \log \frac{P_m}{P_d} + u_t$$

where u_t is an error term.

In the linear form the income and price elasticities of import demand will depend on the levels of these variables, while in the log-linear form the income and price elasticities will be constant and are measured by the coefficients β_1 and β_2, which are read directly from the regression result. There are no clear-cut criteria that can be relied on in choosing a functional form. The choice of the form is usually based on practical considerations and intuition. For the choice between the two forms, Sargan (1964, p. 42) suggests the maximum likelihood ratio $S/Y_G s$, where S and s are the standard errors of the linear and log-linear regressions respectively, and Y_G is the geometric mean of the dependent variable Y_t. If $S/Y_G s$ is less than one we accept the linear-form hypothesis, and vice-versa. In most of our equations the above ratio indicated that the log-linear form was preferred to the linear one.

Finally, the foreign demand functions for Greek exports may be formulated in a way analogous to that used for the country's import demand relationships. Greek exports face a competitive world market and according to economic theory, relative prices and world income are the two major variables to explain the behavior of exports. The export demand functions, in linear and log-linear form, may be written as

$$X = c_0 + c_1 I_w + c_2 \frac{P_x}{P_w} + u_t$$

and

$$\log X = C_0 + C_1 \log I_w + C_2 \log \frac{P_x}{P_w} + u_t$$

where

X = Quantity of the exported group of commodities

I_w = Real income or activity variable of the world

P_x = Export price of Greece

P_w = Prices of related commodities in foreign markets, and

u_t = an error term .

2.2. The Relevance of Least Squares

The estimates of price and income elasticities set out in this study were obtained by least squares regression of quantity of imports on activity variables and relative prices. The least squares method (linear or non-linear according to the error specification) was adopted assuming that each import demand equation is not part of a system of equations in which there is multilateral dependence between the dependent and one or more explanatory variables of the equation. That is, we assume that the disturbances are independent of the explanatory variables and so the estimates are free of simultaneity bias.

However, the application of least squares to import or export demand equations has been criticized on various grounds. A number of empirical studies, appearing in the early post-war years (see Cheng (1959) for a description of these studies), suggested that the international price elasticities are exceedingly low. This suggested in the context of exchange-rate adjustment that a devaluation would tend to worsen rather than improve the trade balance because the sum of the price elasticities of demand for a country's imports and exports might together add up to less than unity. That is, if we assume infinite supply elasticities, the Marshall-

Lerner condition that is necessary for a devaluation to improve the trade balance would not be satisfied. This "elasticity pessimism" suggested that measures other than changes in relative prices might have to be relied upon for purposes of adjusting the balance of trade.

These surprisingly low elasticities have stimulated critical examinations of whether the application of least squares to import or export demand equations and the data available tended to bias the calculated international price elasticities towards zero (Orcutt (1950), Harberger (1953), Machlup (1950)). Orcutt, one of the main critics, gives the following five reasons why the least squares estimates of the foreign trade price elasticities are biased towards zero.

(i) Lack of independence between relative prices and the random deviations (shifts) in the import-demand function (least-squares bias)

(ii) Errors of observations in explanatory variables

(iii) The use of aggregated data may give undue weight to goods with relatively low elasticities

(iv) Short-run elasticities were measured and these tend to be lower than the long-run elasticities

(v) Devaluation elasticities are larger than the estimated short-period elasticities, which reflect adjustment to small price changes.

His arguments have stimulated a number of critical discussions (Prais (1962), Liu (1954), Klein at al (1961), Kemp (1962), Magee (1975)) and it has been shown that under certain conditions the least squares procedure is reasonably applicable. The various sources of bias, except the simultaneity bias, are general methodological difficulties related to any application of econometric methods.

As regards the least squares bias, Orcutt's argument is that the quantity and the price variables satisfy not only the import demand equation but also the supply equation, and the price elasticity estimated by the method of least squares will be biased towards zero. However, according to Klein (1960, p. 871), "international trading relationships pitting a small country's demand or supply against an overwhelming world market may also properly be estimated by the ordinary method of least squares". In these cases, the import price can be treated as predetermined and the least squares bias is expected to be negligible or even zero.

It will be argued that, for most categories of Greek imports, certain sufficient conditions for unbiasedness hold to a satisfactory approximation. Since the quantity and price variables satisfy both demand and supply relations, the least squares estimate of the price elasticity will in principle be a weighted average of a negative demand elasticity and a positive supply elasticity; that is, the estimate of the demand elasticity will be biased towards zero and in some cases the estimate may become positive. The estimated price elasticity will be nearer the demand elasticity the larger are the shifts in the supply schedule relative to those of the demand schedule. But the bias disappears completely if the supply schedule is perfectly elastic and the shifts in the supply curves are independent of the shifts in the demand curve.

Greece does not possess any monopsonistic power in the world market and its imports constitute a very small proportion of the world's output. It follows that the Greek import supply functions must be highly elastic and without serious error they may be assumed to be infinitely elastic. It is also expected that the shifts in foreign supply and Greek import demand will be independent. This is supported by the fact that the shifts in the supply curves of imports are determined largely by the changes in the general price levels in the exporting countries and are therefore quite independent of the shifts in the demand curve of Greek imports. So Greece is a price taker and its import price can be treated as a predetermined variable.

The general price level of Greece may also be considered as predetermined because it is largely determined by such broad factors as aggregate money income, total output, wage levels, and the general economic policy of Greece. Therefore, the relative price (P_m/P_d) entering the import demand equation may be treated as predetermined variable.

The above can also be treated formally as follows. Consider the linearized version of our import demand function for a group of commodities.

$$M = a_0 + a_1 I_G + a_2 \frac{P_m}{P_d} + u$$

where u is an error term with zero mean. Under the assumption of the infinitely elastic supply function, the import price-forming relation may be written as

$$(3) \qquad p_m = \beta_1 \, d_w + \beta_2 \, p_w + v$$

where d_w is the foreign demand, p_w is a general index of prices in the rest of the world and v is an error term with zero mean. Consider now the foreign demand relation

(4) $$d_w = \gamma_1 i_w + \gamma_2 p_m + \gamma_3 p_w + w$$

where i_w is the real income of the rest of the world and w is an error term with expectation zero. Notice that the variables of equations (3) and (4) are expressed as deviations from their respective means (lower-case letters stand for deviations). Substituting for d_w in (3), we have

(5) $$p_m = \beta_1\gamma_1 i_w + \beta_1\gamma_2 p_m + (\beta_1\gamma_3+\beta_2)p_w + \beta_1 w + v$$

Multiplying (5) by u and taking expectations we get

(6) $$E(p_m u) = \frac{\beta_1\gamma_1 E(i_w u) + (\beta_1\gamma_3+\beta_2)E(p_w u) + \beta_1 E(wu) + E(vu)}{1 - \beta_1\gamma_2}$$

Since Greece is too small to affect world income and prices, it is reasonable to consider these variables as exogenous to our model; i.e. to treat $E(i_w u)$ and $E(p_w u)$ as taking zero values. Then (6) becomes

$$E(p_m u) = \frac{\beta_1 E(wu) + E(vu)}{1 - \beta_1\gamma_2}$$

Because of the different stages of economic development among Greece and its important trading countries as well as of their different relative sizes, it is rather expected that the shifts in foreign demand and Greek import demand will be independent; that is $E(wu) = 0$. Moreover, if constant returns to scale prevail in the exporting industries, then the export supply will be infinitely elastic, i.e. $\beta_1=0$ (Kemp (1962), p. 14). Finally, the independence between the foreign supply and Greek import demand implies that $E(vu)=0$. Hence $E(p_m u) = 0$, suggesting that import prices can be considered as predetermined. Since domestic prices are treated as predetermined too, the relative prices can also be treated as such.

Moreover, as it can be observed, the above system is recursive and therefore the least-squares method is legitimate. In particular, expressing the import demand equation as

$$m = a_1 i_G + a_2 p_m + a_3 p_d + u$$

and treating i_g and p_d as exogenous, then with the assumption that in the import price equation (3) $\beta_1 = 0$, the matrix of the coefficients of the endogenous variables is triangular. On the other hand assuming that constant return to scale prevail in the exporting industries, i.e. $\beta_1 = 0$, allows us to treat E(vw) as taking zero value, that is the shifts in foreign demand are independent of the shifts in foreign supply. This in connection with the assumptions mentioned before on the independence among the errors of the import-demand equation of Greece and the errors of the equations (3) and (4), gives a diagonal variance-covariance matrix of the disturbances. Hence the system is recursive and we can apply least-squares on each equation separately without any bias.

We can also show that if the imported group of commodities is quite small in relation to the income of Greece so that fluctuations in the demand for that particular category of imports have a negligible effect on income or economic activity of Greece, then the income variable can be treated as a predetermined variable. This is supported by the fact that the bulk of the variation in income is usually attributed to changes in investment and goverment expenditures and therefore imports are not expected to have any noticeable feed back effects on income. For that reason it appears that the income variable may be approximately treated as a predetermined variable in connection with our import demand equations.

Estimates of the coefficients of the Greek export demand equations are also obtained by the application of the method of least squares. Greece exports a small fraction of total world imports. Her international price system is fully reflected by world market prices, and therefore export prices are assumed to be exogenous variables (Klein at al (1961), p. 130). The rest of the main variables influencing the foreign demand for Greek exports, such as world prices and world activity, are also exogenous. Thus, in the case of a small country like Greece with a small participation in the world trade, Orcutt's difficulties arising from the least squares bias are unimportant and even non-existent.

Orcutt's second point was that when the data contain errors of measurement due to misclassification, falsification, and faulty methods of index-number construction, the effect may be to bias the coefficients towards zero. The consequences from these errors depend on the assumptions made on the relations between the observation errors and the true values (see Johnston (1972), ch. 9, for a detailed treatment). Orcutt's argument is based on the assumption that errors in the quantity (dependent) variable are uncorrelated with the observed price and income variables (Orcutt, 1950, p. 129). Kemp (1962) has shown that if the errors are independent of the true value of the price variables and the errors in the quantity variable are negatively correlated with the errors in the price variable, then the bias with which the price elasticities are estimated is not towards zero but towards unity. Therefore, since usually nothing is known about the true relationship between errors and true values, it is difficult to reach any conclusion.

In his third point Orcutt argued that goods with relatively low elasticities are more heavily weighted in aggregated data than high-elasticity products. The use of aggregative price indexes may thus understate the true price elasticity. In this study we have adopted the one-digit SITC level of aggregation. This was largely dictated by the presentation of the basic data and we believe that the possible aggregation bias in the estimation of the price elasticities has been reduced satisfactorily.

Orcutt's fourth point related to the fact that what was estimated in most studies was a short-run elasticity that would be expected to be lower than the long-run elasticity. Klein at al (1961, p.p. 135-136) argue that it is not possible to generalize about the relation between short-run and long-run elasticities, while Leamer and Stern (1970) describe the assumptions under which a short-run elasticity is meaningful. However, the estimation of short-run and long-run elasticities is related to the dynamic specification of the import demand functions, and this has been discussed in the preceding sections.

Finally, Orcutt's fifth point was that the price elasticity of demand for large price changes will generally be higher than for small price changes observed in the used price series. Thus, extrapolating beyond the limits of sample experience might be dangerous. Although we agree with Orcutt's argument, in our case the sample runs from January 1954 to December 1976 and thus it allows for relatively large price changes. Therefore, extrapolating will be less dangerous and the standard error of forecast will be small.

In conclusion, Orcutt's arguments seem to be overstated and in Prais's (1962, p. 575) words, "His arguments, taken by themselves, should thus be interpreted as being in favor of more careful research, rather than requiring its abandonment". We believe that prerequisite attention has been paid for a more careful research in this study, as far as possible.

CHAPTER V

THE ESTIMATED IMPORT DEMAND EQUATIONS OF GREECE

The preceding chapter dealt with the empirical specification of dynamic models, with the adopted general hypothesis to be tested as well as with the selection of the econometric method to be used for testing that hypothesis. In the present and next chapters an attempt will be made to obtain the preferred specifications and their numerical estimates of the import and export demand functions of Greece, applying the model selection procedures described before.

1. Statistical Data

The data for imports used in this study are time series observations for the period 1954-1976. As regards the choice of the unit of observation, it was decided to use monthly data, instead of annual or quarterly data which have been employed by other researchers. Data before 1954 are not considered, firstly because they are not available on a monthly basis and secondly because during the period 1949-1953 severe import restrictions were imposed by the authorities to check the large deficits in the balance of payments. In April 1953, the Greek currency (drachma) was devaluated by 50 percent (the new official parity was set at drachmae 30 = U.S. $1 and the exchange rate remained unchanged till 1974 when it was decided, due to the instability of U.S. dollar, to let drachma fluctuate according to a group of the most stable foreign currencies). Since then Greece has pursued a liberal import policy to the extent that only a very small proportion of imports requires import licences.

Six major groups of imported commodities are examined in this study: food, raw materials, fuels, chemicals, manufactures, and machinery and transport equipment. This classification coincides with the Standard International Trade Classification (the above groups are the corresponding SITC categories 0, 2, 3, 5, 6 and 7), and was largely dictated by the presentation of the basic data. The omitted SITC categories (1, 4, 8 and 9) are not examined because being unimportant for Greek import trade, they are not listed for the period before 1972. They are, however, included in total imports.

The basic statistical data of Greek imports were taken primarily from the Monthly Bulletin of External Trade Statistics of Greece, published by the National Statistical Service of Greece. Imports are valued on cost, insurance and freight (c.i.f.) basis, and are recorded at the clearance point of the Customs authorities. The trade statistics give volume indices of the Laspeyres type as well as price indices of the Paasche type for the total imports and the major groups of commodities mentioned above. The price indices are unit values obtained as the ratio of the current value of imports to their value at constant prices. Since the indices are given in three separate series based on 1954, 1961 and 1970 prices respectively, continuous series for the sample period were obtained by splicing together the three separate series.

Furthermore, the import prices were adjusted for changes in duties and taxes, because, on the one hand some of the imported commodities are subject to specific tariffs and, on the other hand, all imports from the European Economic Community have been subject to a gradual tariff reduction since Greece's association with that area in November 1962.

The current weighted import price index adjusted for changes in duties and taxes is given by

$$P_a = \frac{\Sigma(p_{in} + t_{in})\, q_{in}}{\Sigma(p_{i0} + t_{i0})\, q_{in}} =$$

$$= \frac{\Sigma p_{in}q_{in} + \Sigma t_{in}q_{in}}{\Sigma p_{i0}q_{in} + \Sigma t_{i0}q_{in}} =$$

$$= \frac{\Sigma p_{in}q_{in}}{\Sigma p_{i0}q_{in}} \cdot \frac{1 + (\Sigma t_{in}q_{in}/\Sigma p_{in}q_{in})}{1 + (\Sigma t_{i0}q_{in}/\Sigma p_{i0}q_{in})} =$$

$$= P \cdot \frac{1 + (\Sigma t_{in}q_{in}/\Sigma p_{in}q_{in})}{1 + (\Sigma t_{i0}q_{in}/\Sigma p_{i0}q_{in})}$$

where, t_{in} and t_{i0} are the duties and taxes per unit of commodity i for the current (n) and base (0) periods respectively, and P is the unadjusted import price index.

Assuming that $\frac{\Sigma t_{i0}q_{in}}{\Sigma p_{i0}q_{in}} \approx \frac{\Sigma t_{i0}q_{i0}}{\Sigma p_{i0}q_{i0}}$, we have

$$P_a = P \cdot \frac{1 + (\Sigma t_{in}q_{in}/\Sigma p_{in}q_{in})}{1 + (\Sigma t_{i0}q_{i0}/\Sigma p_{i0}q_{i0})} =$$

$$= P \cdot \left(\frac{1 + R_n}{1 + R_0} \right)$$

where, R_n and R_0 are the ratios of duties and taxes to the value of imports at the current and base periods respectively.

As it was mentioned in the preceding chapter, economic theory suggests that the main determinants of the import demand for a commodity or a group of commodities are an income or activity variable and the import price relative to domestic prices. In the case of imported commodities which are objects of final consumption real disposable income should be selected as the income variable, while in the case of imported raw materials the level of activity of the consuming industry provides the proper activity variable. However in Greece monthly figures for income are not published and the index of industrial production or its appropriate component has been selected as an income-proxy (e.g. for imports of chemicals the index of chemical production is employed as activity variable). The use of the index of industrial production as a proxy income variable is not unreasonable. In addition to the high intercorrelation between the real national product and the index of industrial production, the latter seems to be more plausible in the case of Greece, where the urban sector is the main source of demand for imported commodities.

To measure the substitutability between imported and domestically produced goods the import price, corrected for changes in duties and taxes, is usually deflated by a price index of domestically produced import competing products. But the limited competitiveness of Greece's production does not justify the application of this technique to the present case. Instead, considering that in a broader sense imports are competitive with all other domestically sold goods, the domestic wholesale price indices are the most appropriate deflators. The price series we used in this study are the general index of wholesale prices and the corresponding indices for food, raw materials and finished goods according to the nature of the imported group of commodities.

Apart from price and activity variables, other factors may also exert an influence on the import demand for a particular group of commodities. In particular, receipts of the country from exports of goods and services were used as a proxy variable for the strigency of controls affecting imports. In other words this variable was introduced into the equations in order to measure the ability of the country to import. Also a time trend was included in the import demand equations to take account of any systematic changes in other influences which have not been introduced explicitly into the equations, as for example changes in tastes, technology and the like. However these variables failed to produce significant results because of the high intercorrelation between them and the index of industrial production.

Finally, with respect to the rest of the data (domestic prices, activity variables, import duties and taxes etc.) the the principal sources are various issues of the Statistical Yearbook of Greece, the Monthly Statistical Bulletin of Greece, the Monthly Statistical Bulletin of Public Finance of Greece (all published by the National Statistical Service of Greece) and the Monthly Bulletin of the Bank of Greece.

2. The Estimated Import Demand Equations of Greece for Major Groups of Commodities

2.1. Import Demand Equations for Food (SITC: Section 0)

Imports of food amounted to 11.6 percent in the value of the total imported goods over the period under consideration. Imported food items, such as meat, dairy products, fish, cereals, animal feeds, coffee and cocoa account for more than 70 percent in the imports of food.

The share of food in the total imports of goods decreased from 16.8 percent in 1954 to 9.4 percent in 1976 (chapter II, table 3). During the same period the composition of imported food has been changed considerably. In particular, the share of meat in the total imports of food increased from 5% in 1954 to 30% in 1976. This increasing import demand for meat is mainly attributable to the rise in the standard of living and to the inadequacy of the domestic supply to satisfy the increasing demand for this product. Imports of fish, which amounted to 11.2% of the total imported food in 1954, reduced to 5.4% in 1976 because of the development of the Greek fishery during the period under review.

The share of cereals in the total imports of food remained more or less unchanged during the period 1954-76 and it averaged about 25 percent. However, the composition of imported cereals has been changed completely. For instance, imports of wheat unmilled which accounted for 23% of the total imported food in 1954 decreased to a negligible proportion in 1976. Since the early sixties, Greece has become self-sufficient in wheat. This was mainly due to improved methods of agricultural production, such as better seeds, use of fertilizers, and introduction of mechanization. On the contrary, the share of maize unmilled, which is used mainly as animal feed, in the imports of food increased from 0.5% in 1954 to 20% in 1976. During the same period the share of imported animal feeds increased from 0.6% to 7.8%. The above increasing demand for imported animal feeds is a result of the development of cattle-raising which in view of the increasing demand for meat, as mentioned before, has been expanded considerably. As a result, imports of live animals which accounted for 11.5% in total imports of food in 1954, reduced to 2.1% in 1976.

In the early sixties three factories were established for the production of sugar from domestically produced sugar-beet. Since then the domestic production of sugar has gradually increased and as a result the share of sugar in the imports of food decreased from 13.6% in 1954 to 1% in 1976. Finally, the imports of dairy products and coffee and cocoa remained almost unchanged and their shares in the imports of food were 12.7% and 9.8% respectively over the period 1954-1976.

This account of the composition of imported food implies that the increasing real income and the rise in the standard of living of the Greek people are the main reasons for the shift of import demand for food to items of higher quality such as meat, dairy products, coffee and cocoa, etc.

Since the majority (about 60%) of imported food items are not further processed in production, that is, they are objects of final consumption, the total index of industrial production, which is more closely related to real disposable income than any other index of industrial activity, has been selected as the income variable. Most of the imported food items, except coffee, cocoa and some other small items, are also produced at home. Therefore, to obtain the relative price variable the price index of imported food, adjusted for import duties and taxes, was deflated by the domestic wholesale price index of food.

For the basic model, Sargan's maximum likelihood ratio gave a value of 1.53 indicating that the log-linear form is preferred to the linear one. The least-squares estimates of the basic equation obtained from seasonally adjusted data are as follows, with standard errors in parentheses:

$$\log(M_f) = \underset{(0.688)}{5.707} - \underset{(0.131)}{0.720} \log(P_f^m/P_f^d) + \underset{(0.034)}{0.473} \log(IP) \tag{1}$$

$n=263$, R.S.S.$=18.970$, $R^2=0.594$, D.W.$=1.714$, $X^2(15)=34.655$

where,

M_f = Index of volume of imported food

P_f^m = Price index of imported food

P_f^d = Domestic wholesale price index of food

IP = Total index of industrial production

The price and income elasticity have the expected signs and they are highly significant. The Durbin-Watson statistic is in the inconclusive range whereas the $X^2(15)$ Box-Pierce test - statistic for a random correlogram gives a value of 34.655 which is significant. Reestimating (1) subject to a simple ℓth order autoregressive error, we obtained a significant autocorrelation parameter for ℓ = 1, 9 and 12. In all three cases the estimated elasticities were very close to the ones of equation (1) and the original equation, autoregressive specifications were accepted against the unrestricted transformed equations - $X^2(2)$: 0.983, 2.088 and 0.921 respectively. This was the reason to consider a more general autoregressive form of order ℓ, which finally was preferred and in likelihood ratio tests the hypothesis that $\ell=9$ was accepted against other alternative values.

The maximum likelihood estimates are:

$$\log(M_f) = \underset{(0.872)}{6.158} - \underset{(0.165)}{0.805} \log(P_f^m/P_f^d) + \underset{(0.050)}{0.460} \log(IP) \tag{2}$$

$$\begin{aligned} u_t = {} & \underset{(0.063)}{0.143}\, u_{t-1} + \underset{(0.063)}{0.022}\, u_{t-2} + \underset{(0.063)}{0.054}\, u_{t-3} + \\ & + \underset{(0.062)}{0.085}\, u_{t-4} - \underset{(0.062)}{0.166}\, u_{t-5} + \underset{(0.062)}{0.097}\, u_{t-6} - \\ & - \underset{(0.063)}{0.077}\, u_{t-7} + \underset{(0.063)}{0.075}\, u_{t-8} + \underset{(0.063)}{0.133}\, u_{t-9} + \varepsilon_t \end{aligned}$$

$n=263$, R.S.S. = 17.394, $X^2(15) \simeq 6.863$.

The test of the non-linear restrictions imposed by the autoregressive error specification (2) in the corresponding unrestricted transformed equation, which contains one up to 9 months lagged values of the variables $\log(M_f)$, $\log(P_f^m/P_f^d)$ and $\log(IP)$, gave a value of 20.687 for the $X^2(18)$ test statistic, which is not significant; hence, the original equation, autoregressive error specification is preferred.

Considering now the "common factor" analysis, the general unrestricted dynamic model has the form

(3) $$a(L)\log(M_f) = \beta_1(L)\log(P_f^m/P_f^d)+\beta_2(L)\log(IP)+\text{constant}$$

where the scalar polynomials in the lag operator L are of degree 13. The least squares estimates of (3) gave an equation with 42 coefficients, most of which are insignificant because of the collinearity between the different lags of the variables (n = 263, R.S.S. = 15.699, R^2 = 0.664, D.W. = 2.019, $X^2(15)$ = 4.656).

The Wald test for a common factor polynomial of degree 13, gives a value of 12.092 for the $X^2(26)$ test statistic which is not significant and this suggests that we should consider a structural equation form without any lags on the variables. In this case, as already mentioned, according to the likelihood ratio tests, the original equation with a 9th order general autoregressive error specification, is the preferred form.

The Wald test for a common factor polynomial of degree 12 gives a value of 3.903 for the $X^2(24)$ test-statistic. The difference between the two X^2 test statistics, i.e. $X^2(26)-X^2(24)$, which is distributed as $X^2(2)$ gives the significant value of 8.189 and therefore in testing for a m-degree common factor polynomial the hypothesis that m=12 is accepted against m=13. This suggests, on the basis of the difference between m=12 and m=13, that we require a structural equation form in which at least some of the variables of equation (1) have one extra lag. The same test was also applied in the simplified version of (3), since the preferred specification (2) showed that the initially extracted set of 13 common roots, contained four zero-ones which shorten the lag length by four for every variable. Reestimating (3) allowing this time the scalar polynomials to be of order 9 and testing for m = 9 we obtained the non-significant value of 15.207 for the $X^2(18)$ test-statistic. But testing for m = 9 against m = 8 we obtained the significant value of 12.515 for the $X^2(2)$ test-statistic.

Experimenting with a structural equation which includes all variables lagged one month, only the lagged dependent variable produced a significant coefficient, and among other forms a simple 5th order autoregressive error process was preferred.

The estimated equation is:

$$(4) \qquad \log(M_f) = \underset{(0.710)}{5.054} + \underset{(0.060)}{0.155} \log(M_{f,t-1}) - \underset{(0.126)}{0.655} \log(P_f^m/P_f^d) + \underset{(0.041)}{0.395} \log(IP)$$

$$u_t = - \underset{(0.061)}{0.133}\, u_{t-5} + \varepsilon_t$$

$n = 263$, R.S.S. $= 18.288$, $X^2(15) = 25.843$

The $X^2(1)$ test statistic on ρ_5 gave a value of 4.653 which is significant compared with a $X^2(1)$, 5% confidence limit of 3.84. The autoregressive error hypothesis is also accepted against the unrestricted transformed regression equation - $X^2(3) = 3.599$. Thus, according to the Wald difference criterion, specification (4) is the preferred form. In terms of residual variance model (2) is preferred to model (4) (s^2 equal to 0.069 and 0.071 respectively). Moreover, equation's (4) Box-Pierce test-statistic exceeds 24.996 which is the 5% confidence limit for a $X^2(15)$ and in fact there are peaks in the residual correlogram at lags 6, 8, and 9. Notice, however, that the long-run elasticities derived from equation (4) are very close to the elasticities given from the static equation (2).

Turning now to a test of the seasonal adjustment of the data, we reestimated equation (2) using unadjusted data. The F-test constructed from the corresponding residual sums of squares gave a value of 4.130 for the $F(11,240)$ test statistic which is highly significant. A significant value was also obtained in equation (4) - $F(11,247) = 3.660$. Finally, in both specifications (2) and (4), the log-linear form was preferred to the linear form according to the likelihood ratio tests.

It appears from the above equations that imports of food are treated as essentials by the Greek consumer. This is not a surprise. The income elasticity of import demand for food is generally expected to be below unity. The increasing real income and the rise in the standard of living are the main reasons that the imported food items although of higher quality (meat, fish, dairy products, coffee and cocoa, etc.) are not treated as luxuries by the Greek consumer. However, since industrial production has grown faster than national income, the income elasticity has been evidently underestimated (during the sample period national income and industrial production increased at an average annual rate of 6.3% and 8.4% respectively).

The price elasticity estimated from equation (2) is highly significant and it amounts to -0.805. It appears, then, that some degree of substitutability exists between imported and home produced food. This was expected since, as already mentioned, most of the imported food items are also produced at home. Moreover, the increasing demand for higher quality goods and the increasing inadequacy of the domestic supply to meet it justify a price elasticity of less than unity.

2.2. Import Demand Equations for Crude Materials (SITC: Section 2)

For the period under review imports of crude materials accounted for about 10 percent of the value of total imported goods. This group includes a variety of commodities, but the bulk of imported crude materials consists of hides and skins, wood (rough and simply worked), pulp and waste paper, wool, synthetic fibres, oil-seeds and fertilizers (the share of these items in the total imports of crude materials averaged at about 73 percent during the period 1954 - 76 and their individual shares were 6%, 22%, 9%, 20%, 7%, 4%, and 5% respectively).

Imported crude materials are used to a large extent as inputs by the industrial sector towards the production of final goods. Therefore, the level of industrial production appears to be the most suitable activity variable. Changes in stocks may also have affected the import demand for crude materials during the sample period. Unfortunately, lack of data limits the possibility of measuring the influence of stock fluctuations. Finally, changes in relative prices, that is, the price index of imported raw materials corrected for duties and taxes relative to the domestic wholesale price level of raw materials, were introduced into the equation.

Sargan's maximum likelihood ratio gave a value of 1.46, for the basic equation, which clearly indicates that the log-linear form is preferred to the linear form.

The least squares estimates of the basic model, from seasonally adjusted data, are as follows:

$$\log(M_{cm}) = \underset{(0.818)}{2.441} - \underset{(0.142)}{0.488} \log(P^m_{cm}/P^d_{rm}) + \underset{(0.039)}{0.923} \log(IP) \qquad (1)$$

$n=263$, $RSS=7.458$, $R^2=0.922$, $D.W.=1.536$, $X^2(15)=113.673$

where,

M_{cm} = Index of volume of imported crude materials

P^m_{cm} = Price index of imported crude materials

P^d_{rm} = Domestic wholesale price level of raw materials

IP = Index of industrial Production

The price and income elasticities have the expected signs and they are significant. The Durbin-Watson statistic is significant revealing a first order serial dependence in the residuals. But higher order autocorrelations are also present in the residuals as is suggested by the highly significant value of the $X^2(15)$ Box-Pierce test - statistic for a random correlogram. In fact there are autocorrelations bigger than twice their standard errors in the residual correlogram at lags 1, 2, 3, 4, 6, 7, 8, 9, 11 and 12. Reestimating equation (1) subject to a simple ℓth order autoregressive error, we obtained, as expected, a significant autocorrelation parameter for the values of ℓ corresponding to the aforementioned autocorrelations in the residual correlogram. In all cases, except for ℓ = 3 and 4, the original equation, autoregressive error specification was accepted against the unrestricted transformed equation.

The above indicate that a general autoregressive error scheme would be more appropriate. So we reestimated (1) allowing this time the error term to follow a general ℓth order autoregression which finally was preferred and in likelihood ratio tests the hypothesis that ℓ = 11 was accepted against other alternative values.

The maximum likelihood estimates are:

(2) $$\log(M_{cm}) = \underset{(1.047)}{1.316} - \underset{(0.179)}{0.264} \log(P^{m}_{cm}/P^{d}_{rm}) + \underset{(0.074)}{0.939} \log IP)$$

$$u_t = \underset{(0.062)}{0.144}\, u_{t-1} + \underset{(0.063)}{0.173}\, u_{t-2} + \underset{(0.064)}{0.090}\, u_{t-3} +$$

$$+ \underset{(0.064)}{0.086}\, u_{t-4} - \underset{(0.064)}{0.084}\, u_{t-5} + \underset{(0.065)}{0.068}\, u_{t-6} +$$

$$+ \underset{(0.065)}{0.038}\, u_{t-7} + \underset{(0.065)}{0.006}\, u_{t-8} + \underset{(0.065)}{0.061}\, u_{t-9} -$$

$$- \underset{(0.064)}{0.038}\, u_{t-10} + \underset{(0.063)}{0.178}\, u_{t-11}$$

$n = 263$, R.S.S. $= 6.149$, $X^2(15) = 8.805$

The test of the non-linear restrictions, imposed in the estimation of the general autoregressive error specification (2), gave a value of 25.249 for the $X^2(22)$ test-statistic, which is not significant; hence, the original equation, autoregressive error specification is preferred.

As can be seen from equation (2), if it were not for the significant coefficient of u_{t-11}, a lower order general autoregressive scheme would be the preferred form. But all the likelihood ratio tests which test jointly the significance of the autoregressive parameters give significant values for the corresponding X^2 test statistics at the 5% size of the test. Only for the hypothesis $\rho_5 = \rho_6 = \ldots. = \rho_{11} = 0$ the likelihood ratio test gives a value of 14.485 for the $X^2(7)$ test statistic which is just significant compared with a $X^2(7)$, 5% confidence limit of 14.067. Estimation subject to a general 4th order autoregressive error scheme yields:

$$\text{(3)} \qquad \log(M_{cm}) = \underset{(1.069)}{1.576} - \underset{(0.186)}{0.322} \log(P^m_{cm}/P^d_{rm}) + \underset{(0.058)}{0.942} \log(IP)$$

$$u_t = \underset{(0.062)}{0.143}\, u_{t-1} + \underset{(0.063)}{0.196}\, u_{t-2} + \underset{(0.063)}{0.101}\, u_{t-3} + \underset{(0.063)}{0.109}\, u_{t-4}$$

$n = 263$, R.S.S. $= 6.497$, $X^2(15) = 21.384$

Turning now to a test of the non-linear restrictions imposed in estimation we find that the original specification, autoregressive error hypothesis (3), is rejected in favor of the lagged structural equation ($X^2(8) = 20.521$). So, the preferred form, after dropping some non-significant variables, is

$$\text{(4)} \qquad \log(M_{cm}) = \underset{(0.778)}{0.992} + \underset{(0.060)}{0.199} \log(M_{cm,t-2}) + \underset{(0.059)}{0.172} \log(M_{cm,t-3}) + \underset{(0.060)}{0.165} \log(M_{cm,t-4}) - \underset{(0.137)}{0.203} \log(P^m_{cm}/P^d_{rm}) + \underset{(0.079)}{0.437} \log(IP)$$

$n=263$, $RSS=6.292$, $R^2=0.934$, $D.W.=1.781$, $X^2(15)=16.682$

So if it were not for the significant coefficient of

u_{t-11} in equation (2), the lagged structural equation (4) with white noise errors would be the preferred form. However, since all the tests are performed at the conventional 5% level of significance, specification (2) should be retained as the preferred form. Notice, however, that in terms of residual variance model (4) provides a slightly better fit giving s^2 0.0245 instead of 0.0247 of model (2).

Following now the "common factor" analysis we begin with the general unrestricted dynamic model

$$(5) \qquad a(L)\ \log(M_{cm}) = \beta_1(L)\log(P^m_{cm}/P^d_{rm}) + \beta_2(L)\log(IP) + + \text{constant}$$

where the scalar polynomials in the lag operator L are of degree 13. The least squares estimates of (5) gave an equation with 42 coefficients from which only the coefficients of $\log(M_{cm})$ lagged 2 and 11 periods as well as the coefficient of log(IP) were significantly different from zero ($n = 263$, R.S.S. $= 5.358$, $R^2 = 0.944$, D.W. $= 1.964$, $X^2(15) = = 5.548$). This suggests that the model can be simplified and the Wald test for a common factor polynomial confirms this view. Testing for a common factor polynomial of degree 13 we obtained a value of 17.106 for the $X^2(26)$ test-statistic which is not significant. So a structural equation form without any lags on the variables should be considered, and the already estimated original equation, autoregressive error specification (2) is the preferred form.

The Wald test for a common factor polynomial of degree 12 gives a value of 4.780 for the $X^2(24)$ test statistic. Thus, the difference between the two X^2 test-statistics, i.e. $X^2(26) - X^2(24)$, gives a value of 12.326 for the $X^2(2)$ test-statistic which is highly significant. This implies that, if the true degree m of the common factor polynomial is less or equal to 12, the difference Wald test is valid and the hypothesis that m = 13 should be rejected in favor of m = 12. Hence, we require a structural equation form in which at least some of the variables of equation (1) have one extra lag. Among all the variables lagged one month, only the lagged dependent variable produced a significant coefficient, and estimating subject to an autoregressive error term, a 12th order general autoregressive form was preferred.

The maximum likelihood estimates are:

$$\log(M_{cm}) = \underset{(0.547)}{0.613} + \underset{(0.109)}{0.718} \log(M_{cm,t-1}) - \tag{6}$$

$$- \underset{(0.097)}{0.122} \log(P^{m}_{cm}/P^{d}_{rm}) + \underset{(0.105)}{0.262} \log(IP)$$

$$u_t = - \underset{(0.124)}{0.613}\, u_{t-1} - \underset{(0.156)}{0.253}\, u_{t-2} - \underset{(0.142)}{0.065}\, u_{t-3} +$$

$$+ \underset{(0.112)}{0.097}\, u_{t-4} + \underset{(0.086)}{0.008}\, u_{t-5} + \underset{(0.078)}{0.089}\, u_{t-6} +$$

$$+ \underset{(0.076)}{0.097}\, u_{t-7} + \underset{(0.076)}{0.070}\, u_{t-8} + \underset{(0.077)}{0.120}\, u_{t-9} +$$

$$+ \underset{(0.077)}{0.006}\, u_{t-10} + \underset{(0.075)}{0.150}\, u_{t-11} + \underset{(0.064)}{0.159}\, u_{t-12}$$

$n = 263$, R.S.S. = 5.998, $X^2(15) = 3.622$

The test of the non-linear restrictions imposed in estimation gives a value of 27.944 for the $X^2(24)$ test-statistic which is not significant; hence, the original equation, autoregressive error specification is preferred. We see that in equation (6), suggested by the Wald difference criterion, the coefficient of $\log(M_{cm,t-1})$ is highly significant revealing a partial adjustment mechanism. It appears, also, that the coefficient of adjustment is rather small (0.282) producing a slow adjustment of current demand to long-run import demand. The long-run income elasticity is almost identical to the one estimated by the static model (2). But there is divergence between the estimated price elasticities (-0.434 and -0.264 respectively). However, in both equations the price coefficient is non-significant. In terms of residual variance model (6) provides a slightly better fit giving s^2 0.0243 instead of 0.0247 of the model (2).

To test the seasonal adjustment of the data we reestimated both equations (2) and (6) using unadjusted data. In the case of model (2) the F-test constructed from the corresponding residual sums of squares gave a value of 2.434 for the $F(11,238)$ test-statistic which is significant. But as far as model (6) is concerned the F-test gave a value of 1.577 for the $F(11,236)$ test statistic which is not significant. Equation (6) estimated from unadjusted data gives:

(7)
$$\log(M_{cm}) = \underset{(0.484)}{0.365} + \underset{(0.081)}{0.750} \log(M_{cm,t-1}) - \underset{(0.086)}{0.079} \log(P^m_{cm}/P^d_{rm}) + \underset{(0.091)}{0.242} \log(IP)$$

$$\begin{aligned} u_t = & - \underset{(0.102)}{0.662}\, u_{t-1} - \underset{(0.135)}{0.315}\, u_{t-2} - \underset{(0.132)}{0.134}\, u_{t-3} + \\ & + \underset{(0.115)}{0.015}\, u_{t-4} - \underset{(0.095)}{0.001}\, u_{t-5} + \underset{(0.084)}{0.122}\, u_{t-6} + \\ & + \underset{(0.079)}{0.171}\, u_{t-7} + \underset{(0.079)}{0.089}\, u_{t-8} + \underset{(0.080)}{0.095}\, u_{t-9} - \\ & - \underset{(0.080)}{0.019}\, u_{t-10} + \underset{(0.077)}{0.097}\, u_{t-11} + \underset{(0.064)}{0.180}\, u_{t-12} \end{aligned}$$

$$n = 263, \ R.S.S. = 6.438, \ X^2(15) = 3.297$$

which is also accepted against the unrestricted transformed equation ($X^2(24) = 36.187$) and its estimated coefficients are similar in value to the ones of equation (6).

The industrial production elasticity estimated by equation (2) is highly significant, while its size slightly below unity is in agreement with a priori expectations. The volume of imported raw materials is expected to change in much the same proportion as the level of industrial production and the above income elasticity confirms this proportionality hypothesis.

The price elasticities estimated from models (2) and (7), although with the right sign, do not differ significantly from zero. This indicates that changes in relative prices exert a small influence on the demand for imported raw materials. This is not unreasonable. The import demand for raw materials is in the nature of derived demand and most imported items in this group are specialized materials with limited substitutability. Moreover, industrialization of the country generates a high demand for raw materials which, since are not produced at home, it is justified to be price inelastic.

2.3. Import Demand Equations for Fuels (SITC: Section 3)

The share of fuels in the value of total imports of goods averaged about 11 percent during the period 1954-1972, and about 19 percent during the period 1973 - 1976 when the prices of crude oil increased considerably. This group includes such items as coal, coke crude petroleum and petroleum products. Greece, since 1959, has started to operate a number of state oil refineries and as a result of that most of the petroleum products are domestically produced by refining of imported crude petroleum. Therefore, crude petroleum only, makes up about 90 percent of the imported fuels.

Since fuels are not produced at home, to obtain the relative price variable the price index of imported fuels, adjusted for import duties and taxes, was deflated by the general index of wholesale prices, and the index of industrial production has been selected as income - proxy.

Apart from the above explanatory variables, the relative progress in industrialization and modernization of the country and the increasing availability of hydroelectric energy may have affected the import demand for fuels during the sample period. However, experiments made with a time trend accounting for the effect of these influences did not produce significant results because of the high intercorrelation between the index of industrial production and the time variable.

According to the likelihood ratio test, performed in the basic form presented below, the log-linear form was preferred to the linear form. However, all the minimization algorithms which are utilized by Hendry's program, for the estimation of general autoregressive schemes of the disturbances, failed to produce acceptable results when we used logarithms of the variables, because of different local minima. Therefore, the linear form was adopted for this group of imported commodities.

Using seasonally adjusted data, the least squares estimates of the basic equation are as follows, with standard errors in parentheses:

$$M_{fu} = \underset{(9.558)}{-26.357} + \underset{(0.053)}{0.053}\,(P^{m}_{fu}/P^{d}_{g}) + \underset{(0.065)}{1.279}\,IP$$

$$n=263,\ RSS= 545131.411,\ R^2=0.597,\ D.W.=1.808, X^2(15)=40.446$$

where,

M_{fu} = Index of volume of imported fuels

P^{m}_{fu} = Price index of imported fuels adjusted for import duties and taxes

P^{d}_{g} = Domestic general index of wholesale prices

IP = Index of industrial production

The price coefficient comes out with the wrong sign, but it is insignificant. As was mentioned before, the price of crude oil increased disproportionately in comparison with domestic prices during the period 1973-1976, and that caused the relative prices to increase and fluctuate at a higher level than the period before 1973. On the other hand, during the period 1973-1976, imports of fuels increased because of other non-price reasons. In particular, two new refineries were established in 1972, whose products (petroleum products refined) were intended mainly for export (see table 3, chapter II).

This implies that an upward shift occured in the import demand equation for fuels, and to take account of this structural change, we introduced into our equation a dummy variable taking the value one for the period 1973-1976 and zero outside it.

The new estimated equation is:

$$(1) \qquad M_{fu} = \underset{(14.988)}{10.183} - \underset{(0.068)}{0.081}\,(P^{m}_{fu}/P^{d}_{g}) + \underset{(0.122)}{0.957}\,IP + \underset{(14.928)}{46.726}\,D$$

$n=263$, $RSS=525260.360$, $R^2=0.611$, $D.W.=1.865$, $X^2(15)=44.848$

where,

$D = 1$ for 1973 - 1976 and 0 elsewhere

The income and the price coefficients have the expected signs, but the price coefficient is still insignificant. The elasticities computed at the point of the sample means are 0.916 and -0.138 respectively and they are in general agreement with a priori expectations. Since fuels are used by both consumers and producers and considering the progress in industrialization, an income elasticity in the neighbourhood of unity is reasonable. Also it is a priori expected imports of fuels to be price inelastic because fuels are not produced at home and are considered as necessities. The significant positive coefficient of the dummy variable indicates an upward shift of the regression plane during the years 1973-76, causing also a significant reduction in the residual sum of squares (RSS). The Durbin-Watson statistic is still non-significant revealing a first order serial independence in the residuals and according to the conventional criteria the estimates are acceptable.

But the $X^2(15)$ Box-Pierce test statistic for a random correlogram gave a value of 44.848 which is highly significant and in fact there is a peak in the residual correlogram at lag 12. So we reestimated the above equation allowing the error term to follow a simple ℓth order autoregressive process and experiments were made for values of ℓ from 1 to 13. As it was expected, according to the likelihood ratio or the asymptotically equivalent t test, only for the 12th order autoregressive process we obtained a significant autocorrelation parameter.

The maximum likelihood estimates are

$$M_{fu} = \underset{(16.048)}{20.898} - \underset{(0.073)}{0.156}\,(P^{m}_{fu}/P^{d}_{g}) + \underset{(0.142)}{0.946}\,IP + \underset{(15.189)}{47.347}\,D \tag{2}$$

$$u_t = \underset{(0.060)}{0.367}\,u_{t-12} + \varepsilon_t$$

$$n = 263,\ R.S.S. = 459019.573,\ X^2(15) = 15.619$$

The $X^2(1)$ test on ρ_{12} gave a value of 35.453 which is highly significant. The standard errors of the estimated coefficients have increased, indicating that were initially underestimated in the presence of autocorrelated errors. Also, all the coefficients came up almost with the same values as in (1) with an exception the price coefficient which increased in absolute value and became significant.

The test of the non-linear restrictions imposed by the autoregressive error specification (2) in the corresponding unrestricted transformed equation, which contains twelve months lagged values of the variables M_{fu}, (P^{m}_{fu}/P^{d}_{g}) and IP, gave a value of 3.105 for the $X^2(2)$ test statistic, which is not significant; hence, the original equation, autoregressive error specification, is preferred.

Due to the collinearity between current and lagged values, the coefficients of the price and income variables came out either with wrong sign or insignificant in the UTE. Only the coefficient of the dependent variable lagged 12 months was significant. Reestimating the UTE after dropping the nonsignificant variables we obtained the following equation:

$$(3) \qquad M_{fu} = \underset{(14.202)}{15.507} + \underset{(0.059)}{0.338}\, M_{fu,t-12} - \underset{(0.064)}{0.122}\, (P^{m}_{fu}/P^{d}_{g}) + \underset{(0.127)}{0.640}\, IP + \underset{(14.033)}{35.033}\, D$$

$n=263$, $RSS=466011.288$, $R^2=0.655$, $D.W.=1.877$, $X^2(15)=14.094$

which, while appearing acceptable on conventional criteria, has a bigger RSS than the autoregressive specification (2). This was expected since equation (2) is preferred according to the likelihood ratio test.

Considering now the "common factor" analysis, the general unrestricted dynamic model has the form

$$(4) \qquad a(L)\, M_{fu} = \beta_1(L)\,(P^{m}_{fu}/P^{d}_{g}) + \beta_2(L)\, IP + \text{constant} + cD$$

where the scalar polynomials in the lag operator L are of order 13. The least squares estimates of (4) gave an equation with 43 coefficients, most of which are insignificant (and with wrong signs) because of the collinearity between the different lags of the variables ($n = 263$, $R.S.S. = 370685.572$, $R^2 = 0.726$, $D.W. = 2.009$, $X^2(15) = 5.271$).

The Wald test for a common factor polynomial of order 13, gives a value of 21.874 for the $X^2(26)$ test statistic which is not significant and this suggests that we should consider a structural equation form without any lags on the variables, and an autoregressive error specification, at least in the first instance, of general 13th order form. The simultaneous estimates of the regression coefficients and the autocorrelation parameters are:

$$(5)\qquad M_{fu} = \underset{(16.101)}{23.684} - \underset{(0.081)}{0.177}\,(P^m_{fu}/P^d_g) + \underset{(0.161)}{0.944}\,IP + \underset{(17.064)}{48.947}\,D$$

$$\begin{aligned} u_t = {} & \underset{(0.064)}{0.045}\,u_{t-1} - \underset{(0.061)}{0.001}u_{t-2} - \underset{(0.062)}{0.025}\,u_{t-3} - \underset{(0.062)}{0.032}\,u_{t-4} - \\ & - \underset{(0.062)}{0.052}\,u_{t-5} - \underset{(0.062)}{0.038}\,u_{t-6} + \underset{(0.062)}{0.065}\,u_{t-7} - \underset{(0.062)}{0.065}\,u_{t-8} - \\ & - \underset{(0.062)}{0.006}\,u_{t-9} + \underset{(0.062)}{0.099}u_{t-10} + \underset{(0.062)}{0.076}u_{t-11} + \underset{(0.062)}{0.351}u_{t-12} - \\ & - \underset{(0.066)}{0.030}\,u_{t-13} \end{aligned}$$

$n = 263$, R.S.S. $= 441767.247$, $X^2(15) = 8.345$

The hypothesis that $\rho_1 = \rho_2 = \cdots = \rho_{11} = \rho_{13} = 0$ in (5) is accepted ($X^2(12) = 10.075$) and so the simple 12th order autoregressive specification is preferred. Moreover, the general autoregressive error hypothesis (5) is rejected in favor of the general unrestricted transformed regression equation (4): $X^2(26) = 46.138$.

The corresponding Wald test for a common factor polynomial of degree 12 gave a value of 4.718 for the $X^2(24)$ test statistic which of course is not significant. But the difference between the two X^2 test statistics, i.e. $X^2(26) - X^2(24)$, which is distributed as $X^2(2)$ gives the significant value of 17.156 and this suggests that we require a structural equation form in which at least some of the variables of equation (1) have one extra lag. However, experiments made with one period lagged values of the variables, and 12th order simple or general autoregressive error specifications, failed to produce significant results (i.e. in all cases the coefficients of the lagged variables were insignificant).

An alternative autoregressive error specification which is also in agreement with the Wald criterion indication for a common polynomial of order 13, is the multiplication of the first order and 12th order quasi-difference operators, namely,

$$(1 - \rho_1 L)\ (1 - \rho_{12} L^{12})\ u_t = \varepsilon_t$$

The estimated equation is:

$$(6) \qquad M_{fu} = \underset{(15.340)}{17.840} - \underset{(0.075)}{0.169}\ (P^m_{fu}/P^d_g) + \underset{(0.137)}{0.993}\ IP + \underset{(14.698)}{46.596}\ D$$

$$(\ 1 - \underset{(0.051)}{0.039}\ L)\ (1 - \underset{(0.061)}{0.355}\ L^{12})\ u_t = \varepsilon_t$$

$$n = 263,\ R.S.S. = 458432.579,\ X^2(15) = 13.841$$

The above specification (6) is accepted against the lagged structural equation which contains one, twelve and thirteen months lagged values of the dependent, price and income variables - $X^2(7) = 7.457$. However, the hypothesis that $\rho_1 = 0$ is accepted ($X^2(1) = 0.337$) and specification (2) is still preferred. This was expected since the most of the reduction in the residual sum of squares of the original structural equation (1) is achieved in both cases (5) and (6) from the highly (and unique) significant value of ρ_{12}.

The non-significant value of ρ_{13} in (5) and the non-significant autocorrelation coefficient ρ_1 in (6) indicate that the existing common polynomial of order 13 contains one root equal to zero and twelve non-zero ones. To verify that we reestimated the general unrestricted equation (4) allowing this time the scalar polynomials in the lag operator L to be of order 12 ($n = 263$, R.S.S. = 372791.986, $R^2 = 0.724$, D.W. = 1.968, $X^2(15) = 5.555$). The likelihood ratio test between the two general unrestricted forms gives the non-significant value of 1.490 for the $X^2(3)$ test-statistic, which shows that the common root of zero shortens the lag length by one for every variable.

The Wald-test for a common factor polynomial of degree 12 gives a value of 22.764 for the $X^2(24)$ test statistic which is not significant, and that indicates again that we don't require any lags in the structural equation. Allowing now the error term to follow a general 12th order autoregression we obtained the following estimates:

$$(7) \qquad M_{fu} = \underset{(16.645)}{24.219} - \underset{(0.085)}{0.189}\,(P^m_{fu}/P^d_g) + \underset{(0.172)}{0.961}\,IP + \underset{(17.208)}{46.836}\,D$$

$$u_t = \underset{(0.061)}{0.038}\,u_{t-1} - \underset{(0.061)}{0.0003}\,u_{t-2} - \underset{(0.062)}{0.023}\,u_{t-3} - \underset{(0.062)}{0.029}\,u_{t-4} -$$

$$- \underset{(0.062)}{0.046}\,u_{t-5} - \underset{(0.062)}{0.036}\,u_{t-6} + \underset{(0.062)}{0.070}\,u_{t-7} - \underset{(0.062)}{0.060}\,u_{t-8} -$$

$$- \underset{(0.062)}{0.0005}u_{t-9} + \underset{(0.062)}{0.105}u_{t-10} + \underset{(0.062)}{0.079}u_{t-11} + \underset{(0.062)}{0.354}\,u_{t-12}$$

$$n = 263,\ R.S.S. = 442125.244,\ X^2(15) = 9.134$$

Here again, the hypothesis that $\rho_1 = \ldots = \rho_{11} = 0$ is accepted ($X^2(11) = 9.862$) and the simple 12th order autoregressive error specification is preferred. Also, specification (7) is rejected in favor of the unrestricted transformed equation ($X^2(24) = 44.861$).

The above empirical analysis indicates that the overall dynamics in the import demand equation for fuels is of 12th order. As regards the systematic dynamics we see that this is due to the significant coefficients of the 12 month lagged variables and mainly to the highly significant value of the coefficient of the lagged dependent variable. The F-test on the coefficients of the variables lagged one up to 11 months gives a value of $F(33,212) = 1.393$ which is not significant and the unrestricted transformed equation which contains only the current and 12 periods lagged values of all the variables is accepted against the general unrestricted equation (4) (the corresponding $X^2(33)$ test-statistic gives a value of 51.618 which is not significant at a 2% size of the test). This is in agreement with the fact that the simple 12th order autoregression form is preferred to the general scheme (7). But, as already mentioned, the non-linear restrictions imposed in the estimation of the simple 12th order autoregressive form are accepted and therefore equation (2) is our data generation process. This is in agreement with the Wald criteria of the common factor analysis which suggested that we require a structural equation without any lags on the variables and with an autoregressive process in the disturbances. So the overall dynamics in the equation is expressed only through error dynamics.

According to the likelihood ratio test, performed in the preferred equation (2), the log-linear form is preferred to the linear form. However, a straightforward shift from the linear form to the log-linear one is not possible, because in the log-linear version of (2) the autoregressive error hypothesis is rejected in favor of the unrestricted transformed equation ($X^2(2) = 17.324$). Moreover, the log-linear form of the basic equation possesses also a significant 3rd order autocorrelation in the disturbances which is also rejected in favor of the unrestricted transformed equation ($X^2(2) = 33.750$). Thus the preferred form when we use logarithms of the variables is, after dropping the non-significant variables,

$$\log(M_{fu}) = \underset{(0.910)}{2.423} + \underset{(0.060)}{0.201} \log(M_{fu,t-3}) + \underset{(0.060)}{0.189} \log(M_{fu,t-12}) -$$

$$- \underset{(0.127)}{0.330} \log(P^m_{fu}/P^d_g) + \underset{(0.119)}{0.375} \log(IP_{t-3}) + \underset{(0.150)}{0.586} D$$

$n=263$, $RSS=56.019$, $R^2=0.662$, $D.W.=1.816$, $X^2(15)=14.324$

which reveals a stronger price effect than the (preferred) linear form. But since a complete empirical analysis using logarithms of the variables is not possible, as it was mentioned in the beginning of the section, we believe that the preferred form based on the levels of the variables should be retained.

Turning now to a test of the seasonal adjustment of the data, we reestimated equation (2) using unadjusted data. The F-test constructed from the corresponding residual sums of squares gave a value of 4.191 for the $F(11,247)$ test statistic which is highly significant. The maximum likelihood estimates of equation (2) including eleven seasonal dummies were as follows:

$$M_{fu} = \underset{(21.440)}{129.857} - \underset{(0.075)}{0.156} (P^m_{fu}/P^d_g) + \underset{(0.145)}{0.946} IP + \underset{(15.509)}{47.418} D -$$

$$- \underset{(20.903)}{127.696} Q_1 - \underset{(20.624)}{132.829} Q_2 - \underset{(20.621)}{115.986} Q_3 - \underset{(20.615)}{126.319} Q_4 -$$

$$- \underset{(20.587)}{103.062} Q_5 - \underset{(20.605)}{121.796} Q_6 - \underset{(20.626)}{133.030} Q_7 - \underset{(20.609)}{140.572} Q_8 -$$

$$- \underset{(20.584)}{118.941} Q_9 - \underset{(20.576)}{86.644} Q_{10} - \underset{(20.610)}{98.921} Q_{11}$$

$$u_t = \underset{(0.061)}{0.369} u_{t-12} + \varepsilon_t$$

$n = 263$, $R.S.S. = 457972.290$, $X^2(15) = 15.783$

There are small differences between the above estimated coefficients and the corresponding ones of equation (2), due to the fact that seasonal adjustment in the latter case is based on the 276 observations of the sample, whereas in the first case the coefficients of Q's are estimated from the last 263 observations of our sample.

The income and price elasticities computed from equation (2), at the point of the sample means are 0.906 and -0.264 respectively. As it was mentioned before, an income elasticity in the neighbourhood of unity is acceptable. But the price elasticity is rather high for such a group of commodities as fuels which a priori are expected to be price inelastic and indicates a noticeable influence of relative prices on the import demand for fuels. However, since 1973 when the oil crisis began, Greek consumers have been facing a number of sucessive increases in the prices of fuels and their reaction to the price changes became more sensitive. Reestimating equation (2) for the period before 1973 we obtained the following estimates:

$$(8) \qquad M_{fu} = \underset{(23.341)}{25.563} - \underset{(0.109)}{0.200}\,(P^{m}_{fu}/P^{d}_{g}) + \underset{(0.183)}{0.963}\,IP$$

$$u_t = \underset{(0.070)}{0.472}\,u_{t-12} + \varepsilon_t$$

$$n = 216,\ R.S.S. = 340726.063,\ X^2(15) = 20.746$$

Here again equation (8) is accepted against the unrestricted transformed equation and the $X^2(1)$ test statistic on ρ_{12} gives the highly significant value of 38.881. The price coefficient is close to the previous one and it is also significant at 7% size of the test suggesting that even with a stable regime in the prices of fuels a price effect exists. Taking also into consideration that about 50% of the imported fuels is used for transportation and heating (the other 50% is used in the industry and for the production of electric energy) the above estimated price elasticity is not unreasonable and suggests that it could be used as an excuse for import controls operating through tariffs.

2.4. Import Demand Equations for Chemicals (SITC: Section 5)

Imports of chemicals include a wide variety of products ranging from raw chemical materials, utilized by the chemical industry, to chemical manufactures such as pharmaceutical products, manufactured fertilizers, inks, and perfumes and other cosmetics in general. This group accounted for about 10 percent in the value of the total imported goods for the period 1954 - 1976.

Since more than one-half of the imported chemicals are used as inputs by the chemical industry, the index of chemical production has been selected to represent the activity variable entering the import demand equation for chemicals. To obtain the relative price variable the price index of imported chemicals corrected for changes in duties and taxes was deflated by the domestic general index of wholesale prices.

For the basic model, Sargan's maximum likelihood ratio gave a value of 1.20 indicating that the log-linear form is preferred to the linear form. The basic equation estimated from seasonally adjusted data, is:

$$(1) \qquad \log(M_{ch}) = \underset{(0.527)}{8.452} - \underset{(0.098)}{1.495} \log(P^m_{ch}/P^d_g) + \underset{(0.020)}{0.676} \log(ICP)$$

$$n=263, \; RSS=8.142, \; R^2=0.944, \; D.W.=1.489, \; X^2(15) = 77.678$$

where,

M_{ch} = Index of volume of imported chemicals

P^m_{ch} = Price index of imported chemicals

P^d_g = Domestic general index of wholesale prices

ICP = Index of chemical production

It appears from the above estimated equation that the activity and price regression coefficients have the expected signs and they are highly significant. The Durbin - Watson statistic is significant revealing a first order serial dependence in the residuals and the $X^2(15)$ Box - Pierce test-statistic gives the highly significant value of 77.678. In fact there are peaks in the residual correlogram at lags 1, 6, 7, 10, 11, 12 and 13. Taking these values for ℓ, a significant autocorrelation parameter was obtained when equation (1) was estimated subject to a simple ℓth order autoregressive process. Also, in all cases, except for $\ell = 10$, the original equation, autoregressive error specification was rejected in favor of the unrestricted transformed equation. But a general ℓth order autoregressive scheme in the errors seems to be more appropriate, which finally was preferred, and in likelihood ratio tests the hypothesis that $\ell = 12$ was accepted against other values of ℓ.

The maximum likelihood estimates are:

$$(2) \qquad \log(M_{ch}) = \underset{(0.598)}{7.258} - \underset{(0.115)}{1.237} \log(P^m_{ch}/P^d_g) + \underset{(0.043)}{0.671} \log(ICP)$$

$$u_t = \underset{(0.062)}{0.205}\, u_{t-1} + \underset{(0.063)}{0.007}\, u_{t-2} + \underset{(0.061)}{0.065}\, u_{t-3} -$$

$$- \underset{(0.059)}{0.042}\, u_{t-4} + \underset{(0.060)}{0.051}\, u_{t-5} - \underset{(0.059)}{0.050}\, u_{t-6} -$$

$$- \underset{(0.060)}{0.070}\, u_{t-7} + \underset{(0.059)}{0.025}\, u_{t-8} + \underset{(0.059)}{0.065}\, u_{t-9} +$$

$$+ \underset{(0.059)}{0.144}\, u_{t-10} + \underset{(0.059)}{0.047}\, u_{t-11} + \underset{(0.058)}{0.204}\, u_{t-12}$$

$n = 263$, R.S.S. $= 6.661$, $X^2(15) = 7.109$.

The highly significant values of the first and twelve order autoregressive parameters ρ_1 and ρ_{12} suggest that the factorised error specification

$$(1 - \rho_1 L)\ (1 - \rho_{12} L^{12}) u_t = \varepsilon_t$$

might be more appropriate. The corresponding unrestricted transformed equation which contains the current and one, twelve and thirteen months lagged values of all the variables, was accepted against the general unrestricted dynamic model, in a F-test ($F(30,210) = 1.285$). Moreover, the Wald test on the validity of the non-linear restrictions (described in chapter IV) imposed by the factorised error specification, gave a value of 7.830 for the $X^2(7)$ test-statistic which is not significant. However, all the minimization algorithms which are utilised by RALS failed to produce acceptable results because of different local minima, and so the above error specification was not considered.

Turning now to a test of the non-linear restrictions imposed in estimation we find that the original specification, autoregressive error hypothesis (2), is rejected in favor of the unrestricted transformed equation ($X^2(24) = 45.041$). This equation, containing one up to twelve months lagged values of all three variables and an independent error term, achieves a substantially lower residual sum of squares than the restricted form given above, and so the preferred form, after dropping the non-significant variables, is:

$$(3)\qquad \log(M_{ch}) = \underset{(0.736)}{7.503} + \underset{(0.053)}{0.170} \log(M_{ch,t-1}) + \underset{(0.046)}{0.164} \log(M_{ch,t-12}) -$$

$$- \underset{(0.112)}{1.102} \log(P^m_{ch}/P^d_g) - \underset{(0.106)}{0.233} \log(P^m_{ch}/P^d_g)_{t-5} +$$

$$+ \underset{(0.048)}{0.393} \log(ICP)_{t-4}$$

$n=263$, $RSS=6.723$, $R^2=0.954$, $D.W.=1.869$, $X^2(15)=16.671$.

Although the deflator is not the proper one, since due to data limitations a price index of domestically produced chemicals is not available, it is worthy to note that the estimated long-run price elasticity of -2.005 in equation (3) reveals a noticeable substitution effect.

Before we consider the "common factor" analysis, we should mention that the seasonal adjustment of the data was not required at any stage of the above empirical analysis. In particular, for models (1), (2), (3) and for the general unrestricted dynamic model, the corresponding F - tests gave the values of 1.318, 0.600, 1.538 and 1.704 for the $F(11,249)$, $F(11,237)$, $F(11,246)$ and $F(11,213)$ test - statistics, respectively, which are not significant. Thus, the above dynamic specification procedure was repeated using this time unadjusted data. Here again, among other forms, a general 12th order autoregressive process was preferred.

The estimated equation is :

$$(4) \qquad \log(M_{ch}) = \underset{(0.591)}{7.439} - \underset{(0.113)}{1.277} \log(P^m_{ch}/P^d_g) + \underset{(0.039)}{0.673} \log(ICP)$$

$$u_t = \underset{(0.062)}{0.200}\, u_{t-1} + \underset{(0.063)}{0.005}\, u_{t-2} + \underset{(0.060)}{0.065}\, u_{t-3} -$$

$$- \underset{(0.059)}{0.045}\, u_{t-4} + \underset{(0.059)}{0.043}\, u_{t-5} - \underset{(0.059)}{0.061}\, u_{t-6} -$$

$$- \underset{(0.059)}{0.079}\, u_{t-7} + \underset{(0.059)}{0.020}\, u_{t-8} + \underset{(0.058)}{0.060}\, u_{t-9} +$$

$$+ \underset{(0.058)}{0.139}\, u_{t-10} + \underset{(0.059)}{0.048}\, u_{t-11} + \underset{(0.058)}{0.228}\, u_{t-12}$$

$n = 263$, R.S.S. $= 6.846$, $X^2(15) = 7.465$.

The structural coefficients and autocorrelation parameters estimated above are almost identical to the ones estimated by equation (2). Here again, in an attempt to estimate the factorised error specification we failed to obtain acceptable results though, both the corresponding unrestricted transformed equation and the validity of the non-linear restrictions were accepted ($F(30,221) = 1.005$ and $X^2(7) = 7.273$).

The test of the non-linear restrictions imposed in estimation gave a value of 30.080 for the $X^2(24)$ test-statistic which is not significant. Hence, for unadjusted data, the original equation, autoregressive error specification (4), is the preferred form.

Considering now the "common factor" analysis, the estimated general unrestricted dynamic model has the form

$$(5) \qquad a(L)\log(M_{ch}) = \beta_1(L)\log(P^m_{ch}/P^d_g) + \beta_2(L)\log(ICP) + \\ + \text{constant}$$

where the scalar polynomials in the lag operator L are of order 13 ($n = 263$, R.S.S. $= 6.090$, $R^2 = 0.959$, D.W. $= 2.016$, $X^2(15) = 9.292$). The Wald test for a common factor polynomial of degree 13, gives a value of 16.347 for the $X^2(26)$ test - statistic which is not significant. This suggests that a structural equation form without any lags on the variables and a general autoregressive error specification, i.e. model (4), should be considered as the data generation process.

The Wald test for a common factor polynomial of degree 12 gives the (non-significant) value of 1.589 for the $X^2(24)$ test statistic. Thus, the Wald difference criterion is equal to 14.758 and exceeds 5.99 which is the 5% confidence limit for a $X^2(2)$, suggesting that we require a structural equation form in which some of the variables of equation (1) have one extra lag. Reestimating the basic equation including this time all the variables lagged one month, only the lagged dependent variable produced a significant coefficient and on estimation subject to an autoregressive error term, a simple 12th order autoregressive form was preferred.

The estimated equation is

$$(6) \qquad \log(M_{ch}) = \underset{(0.605)}{6.797} + \underset{(0.053)}{0.187} \log(M_{ch,t-1}) - \underset{(0.111)}{1.196} \log(P^m_{ch}/P^d_g) + \underset{(0.045)}{0.546} \log(ICP)$$

$$u_t = \underset{(0.057)}{0.261}\, u_{t-12} + \varepsilon_t$$

$n = 263$, R.S.S. $= 7.379$, $X^2(15) = 22.757$

which according to the Wald difference criterion is the preferred form. The $X^2(1)$ test statistic on ρ_{12} gives a value of 18.913 which is highly significant. But the test of the non-linear restrictions imposed by the autoregressive error specification (6) gives a value of 10.685 for the $X^2(3)$ test statistic which is also significant. Hence, the original equation, autoregressive error specification, is rejected in favor of the more complicated lagged structural equation which, after dropping the non-significant variables, gives:

$$(7) \qquad \log(M_{ch}) = \underset{(0.603)}{6.465} + \underset{(0.052)}{0.168} \log(M_{ch,t-1}) + \underset{(0.045)}{0.221} \log(M_{ch,t-12}) - \underset{(0.110)}{1.146} \log(P^m_{ch}/P^d_g) + \underset{(0.048)}{0.369} \log(ICP)$$

$n=263$, $RSS=7.252$, $R^2=0.951$, $D.W.=1.886$, $X^2(15)=19.174$.

We see that the introduction of one period lagged values of the variables into the structural equation, as suggested by the Wald difference criterion, lead us to accept a more general lagged structural equation with an independent error term as the preferred form. We should mention that this is in agreement with our findings in the first stage of the empirical analysis where we considered only simple autoregressive forms in the residuals. In particular, when the original static model (1) was reestimated subject to a simple ℓth order autoregressive error process, in both cases for $\ell = 1$ and 12 the original equation, autoregressive error hypothesis was rejected in favor of the unrestricted transformed equation ($X^2(2)$ had the value of 8.586 and 10.359 respectively) revealing that autocorrelation derives from misspecified dynamics. However, as already mentioned, in likelihood ratio tests a general 12th order autoregressive form was preferred and therefore model (4) should be retained as the preferred specification. Moreover, in terms of residual variance equation (4) provides a slightly better fit giving s^2 0.0276 instead of 0.0286 and 0.0281 of the models (6) and (7) respectively. Finally, in all specifications , i.e. (4), (6) and (7), the log-linear form was preferred to the linear form according to the likelihood ratio test.

The above empirical analysis indicates that the overall dynamics in the import demand equation for chemicals is of 12th order. When 13th order general autoregressive scheme was estimated the hypothesis that $\rho_{13} = 0$ was accepted ($X^2(1) = 0.812$) indicating that the existing common polynomial of degree 13 contains one root equal to zero which shortens the lag length by one for every variable in the general unrestricted dynamic model (5). This was confirmed by a F-test on the coefficients of the variables lagged 13 months, which gave a value of 0.211 for the $F(3,221)$ test-statistic which is not significant. As regards the systematic dynamics we see that this is due to the significant coefficients of the one and twelve months lagged variables and particularly to the lagged values of the dependent variable (the F-test on the coefficients of the variables lagged two up to 11 months gives a value of $F(30,224) = 1.020$ which is not significant). This reflects the fact that in the general autoregressive error form (4), in addition to ρ_{10}, only ρ_1 and ρ_{12} are significant. But, since the non-linear restrictions imposed in the estimation of the general 12th order autoregressive form are accepted, the overall dynamics in the equation is expressed only through error dynamics.

The estimated activity elasticity of 0.673 in equation (4) shows that imported chemicals are treated as essentials by the domestic chemical industry. This activity elasticity is rather low for such a group of commodities, the majority of which consists of raw chemical materials utilised by the chemical industry, and which are expected to change in much the same proportion as the level of chemical production. However, during the sixties several large plants were constructed producing raw chemical materials, such as polysterene, phosphates, aluminium oxide, basic chemicals, petrochemicals, ammonia etc. Therefore, the below unity activity elasticity reflects the gradual shift of the demand from imported raw chemical materials to domestically produced ones.

On the other hand, an activity elasticity below unity shows that the growth rate of demand for imported chemical manufactures is less than the growth rate of demand for domestically produced chemicals. But as far as it concerns the import demand for chemical manufactures, which amount to about 40% of the total imported chemicals, an activity variable which is a closer approximation to real income should be used as an income-proxy. In Greece, the chemical industry has advanced more rapidly than total industry (during the sample period total industrial production and chemical production increased at an average annual rate of 8.4% and 11.6% respectively) and therefore, with respect to the import demand for chemical manufactures, the income elasticity has been underestimated. Reestimating equation (4) using this time the index of total industrial production as an activity variable, we obtained an income elasticity of 1.000[1] and taking into consideration that industrial production has grown faster than national income, the income elasticity of imported chemicals is greater than unity.

The price elasticity estimated by equation (4) is -1.277 whereas the long-run price elasticities obtained from models (6) and (7) are -1.471 and -1.878 respectively, all being highly significant. These above unity price elasticities

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1 Here again a 12th order general autoregressive error form was preferred and the autoregressive error hypothesis was accepted against the unrestricted transformed equation ($X^2(24) = 20.470$).

indicate that there are quite important substitution possibilities between imported and home produced chemicals. This would be expected since, due to the rapid growth of the domestic chemical industry, most of the imported chemicals, and particularly chemical manufactures, are also produced at home.

2.5. Import Demand Equations for Manufactured Goods (SITC: Section 6)

The share of imported manufactures in the value of the total imported goods decreased from 23.8% in 1954 to 15.3% in 1976 and averaged 17.6% over the period 1954 - 1976. The bulk of imported manufactures consists of metals, textiles and clothing, and paper whose shares in the total imports of manufactures are about 50, 19 and 14 percent respectively during the period under review. The remainder of the group includes mainly, rubber, leather, wood and glass manufactures which during the above period amounted to about 5, 4, 3 and 2 percent respectively of the total imported manufactures.

We deal here with a heterogeneous group of commodities, which are destined for many different end uses. Imported manufactures include commodities which are objects of final consumption as well as commodities which are used as inputs in consumer and producer goods industries. Therefore, the index of industrial production appears to be, in this group of imports, the most suitable income variable.

Greece, according to the pattern of its industrial production, lies somewhere between the early stage of industrialization characterised by the development of textile and other light consumer industries, and the latter stages of industrialization in which the production of capital goods, based on modern technology, is introduced. Consequently, it is expected that some degree of substitutability must exist between domestically produced and imported manufactures. To measure this substitutability, the price index of imported manufactures corrected for changes in duties and taxes relative to the domestic wholesale price index of finished manufactures was introduced into the equation.

Sargan's likelihood ratio test, performed in the basic model, gave a value of 1.39 indicating that the log-linear form is preferred to the linear form. Using seasonally adjusted data, the least squares estimates of the basic equation are as follows:

$$(1) \qquad \log(M_{mf}) = \underset{(0.440)}{5.243} - \underset{(0.095)}{0.958} \log(P^{m}_{mf}/P^{d}_{mf}) + \underset{(0.015)}{0.805} \log(IP)$$

$$n=263, \ RSS=4.436, \ R^2=0.930, \ D.W.=1.564, \ X^2(15)=80.857$$

where,

M_{mf} = Index of volume of imported manufactures

P^{m}_{mf} = Price index of imported manufactures

P^{d}_{mf} = Domestic wholesale price level of finished manufactures

IP = Index of industrial production

The activity and price elasticities obtained from the above estimated equation are highly significant and they have the expected signs. Both Durbin-Watson and $X^2(15)$ Box-Pierce test statistics are significant indicating that first order as well as higher order autocorrelations are present in the residuals. In fact, apart from lags 9, 11, 12, 14 and 15, there are autocorrelations bigger than twice their standard errors in the residual correlogram at all other lags. Reestimating equation (1) subject to a simple ℓth order autoregressive error, we obtained, as it was expected, a significant autocorrelation parameter for the values of ℓ corresponding to the aforementioned autocorrelations in the residual correlogram. In all cases the estimated elasticities were very close to the ones of equation (1) and the original equation, autoregressive error specification was accepted against the unrestricted transformed equation. Also, in all the aforementioned equations, the $X^2(15)$ Box-Pierce test-statistic for a random correlogram gave a significant value.

The above indicate that a general autoregressive error scheme should be more appropriate. So we reestimated (1) allowing this time the error term to follow a general ℓth order autoregression which finally was preferred and in likelihood ratio tests the hypothesis that $\ell = 4$ was accepted against other alternative values.

The maximum likelihood estimates are:

(2) $$\log(M_{mf}) = \underset{(0.666)}{5.203} - \underset{(0.143)}{0.937} \log(P^m_{mf}/P^d_{mf}) + \underset{(0.029)}{0.793} \log(IP)$$

$$u_t = \underset{(0.062)}{0.148}\, u_{t-1} + \underset{(0.062)}{0.099}\, u_{t-2} + \underset{(0.062)}{0.133}\, u_{t-3} + \underset{(0.062)}{0.132}\, u_{t-4}$$

$n = 263$, R.S.S. $= 3.975$, $X^2(15) = 14.197$

The $X^2(4)$ test-statistic on the autocorrelation parameters gave a value of 28.827 which is highly significant. The standard errors of the estimated coefficients have increased , indicating that were initially underestimated in the presence of autocorrelated errors, whereas the estimated coefficients decreased in absolute value.

The test of the non-linear restrictions imposed by the autoregressive error specification (2) in the corresponding unrestricted transformed equation, which contains one up to four months lagged values of all the variables, gave a value of 11.296 for the $X^2(8)$ test-statistic, which is not significant; hence, the original equation, autoregressive error specification, is preferred.

Considering now the "common factor" analysis we begin with the general unrestricted dynamic model,

$$(3) \qquad a(L)\log(M_{mf}) = \beta_1(L)\log(P^m_{mf}/P^d_{mf}) + \beta_2(L)\log(IP) + \\ + \text{constant}$$

where the scalar polynomials in the lag operator L are of degree 13. The least squares estimates of (3) gave an equation with 42 coefficients from which only the coefficients of $\log(M_{mf})$ lagged four months as well as the coefficients of $\log(P^m_{mf}/P^d_{mf})$ and $\log(IP)$ were significantly different from zero (n = 263, R.S.S. = 3.532, R^2 = 0.944, D.W. = 1.994, $X^2(15)$ = 2.129). This suggests that the model can be simplified and the Wald test for a common factor polynomial confirms this view. Testing for a common factor polynomial of degree 13 we obtained a value of 10.060 for the $X^2(26)$ test-statistic which is not significant. This suggests that a structural equation form without any lags on the variables, and a general autoregressive specification, i.e. model (2), should be considered as the preferred form.

The difference between the two Wald criteria which test for a common factor polynomial of degree 13 and 12 respectively, gives a value of 7.120 for the $X^2(2)$ test-statistic which is significant. Thus on the basis of the Wald difference criterion the hypothesis that a 13th degree common factor polynomial exists should be rejected in favor of a common factor polynomial of degree 12. This suggests that we require a structural equation form in which at least some of the variables of equation (1) have one extra lag. Among all the variables lagged one month, only the lagged dependent variable produced a significant coefficient, and on estimation subject to an autoregressive error term, a 3rd order simple autoregressive form was preferred.

The estimated equation is:

$$(4) \qquad \log(M_{mf}) = \underset{(0.543)}{4.522} + \underset{(0.057)}{0.160} \log(M_{mf,t-1}) - \underset{(0.113)}{0.829} \log(P^m_{mf}/P^d_{mf}) + \underset{(0.049)}{0.674} \log(IP)$$

$$u_t = \underset{(0.062)}{0.159} u_{t-3} + \varepsilon_t$$

$$n = 263, \; R.S.S. = 4.130, \; X^2(15) = 24.613$$

which according to the Wald difference criterion is the preferred form. But, in a likelihood ratio test specification (4) is rejected in favor of the more complicated lagged structural equation ($X^2(3) = 11.485$), which, after dropping the non-significant variables, gives:

$$(5) \quad \log(M_{mf}) = \underset{(0.526)}{3.564} + \underset{(0.059)}{0.117} \log(M_{mf,t-1}) + $$

$$+ \underset{(0.057)}{0.142} \log(M_{mf,t-3}) + \underset{(0.056)}{0.128} \log(M_{mf,t-4}) -$$

$$- \underset{(0.106)}{0.661} \log(P^m_{mf}/P^d_{mf}) + \underset{(0.061)}{0.490} \log(IP)$$

$n=263$, $RSS=3.988$, $R^2=0.937$, $D.W.=1.993$, $X^2(15)=19.474$.

Thus, the suggestion from the Wald difference criterion to introduce a first order systematic dynamics into the structural equation, leads us to accept the more general lagged structural equation (5) with an independent error term, as the preferred form.

The above preferred specification (2) shows that the initially extracted set of 13 common roots contains nine zero-ones which, in the general unrestricted dynamic model (3), shorten the lag length by nine for every variable. This was confirmed in a F-test on the coefficients of the variables lagged 5 up to 13 months, which gave a value of 0.609 for the $F(27,210)$ test-statistic which is not significant. Reestimating (3) allowing this time the scalar polynomials to be of order 4 and testing for a common factor polynomial of degree 4, we obtained the non-significant value of 6.755 for the $X^2(8)$ test-statistic. Also, the difference between the two Wald criteria which test for a common factor polynomial of degree 4 and 3 respectively, gives a value of 5.406 for the $X^2(2)$ test-statistic which is not significant. Thus, both the Wald and the difference Wald criteria, applied in the simplified version of (3), suggest that we require a structural equation form without any lags on the variables and the original equation with a general autoregressive error specification, i.e. model (2), is the preferred form.

Turning now to a test of the seasonal adjustment of the data, we reestimated equation (2) using unadjusted data. The F-test constructed from the corresponding residual sums of squares gave a value of 3.450 for the $F(11,245)$ test statistic which is highly significant. Finally, according to Sargan's likelihood ratio test, performed in the preferred equation (2), the log-linear form is preferred to the linear form.

The activity elasticity estimated by equation (2) is highly significant, while its size (0.793) below unity is not unreasonable. Most of the imported manufactures are semi-finished goods which are used to a large extent as inputs by the industrial sector toward the production of final goods. These final goods as well as the imported manufactures which are objects of final consumption are not of the luxury type. However, since, as already mentioned, the industrial production has grown faster than national income, we would place the income elasticity of imported manufactures in the neighborhood of unity which is reasonable.

The price elasticity estimated from equation (2) is highly significant and it amounts to -0.937. This indicates that a noticeable substitutability exists between imported and home produced manufactures. This was expected since, as mentioned in the beginning of the section, many of the imported manufactured goods are also produced at home. However, the foreign-made manufactured goods imported in Greece are usually more sophisticated and of greater variety than the home produced and this justifies a price elasticity of slightly less than unity.

2.6. Import Demand Equations for Machinery and Transport Equipment Excluding Ships[1] (SITC: Section 7)

The share of imported machinery and transport equipment, the largest group of Greek imports, in the value of the total imported goods increased from 17.4% in 1954 to 35.4% in 1972, and declined ever since, to reach 27.7% in 1976, whereas it averaged to 29% over the whole sample period. Capital goods make up more than 70 percent of this group, while the rest consists mainly of private passenger cars and domestic electrical equipment.

Imported capital goods are destined to replace and expand the existing capital stock. Therefore, private fixed investment in plant and equipment seems to be the appropriate activity variable for this group of imports. Because of data limitations, however, the index of industrial production has been selected instead as the activity variable. But the selection of the index of industrial production as a proxy variable is not unreasonable. Most of the imported machinery

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[1] Following traditional practice, shipping transactions, whose contribution in the relevant import class was 40% during the sample period, are not examined since they are nothing more than transfers of Greek-owned ships from a flag of convenience to the Greek flag. Moreover, it is not certain whether or not ships would remain under the Greek flag once they came under it.

is used by the industrial sector and private fixed investment is highly correlated with industrialization, that is, with the level of industrial production.

The bulk of the imported commodities in this category, except some domestic electrical equipment, is not produced at home. Thus, to obtain the relative price variable, the price index of imported machinery and transport equipment corrected for duties and taxes was deflated by the domestic general index of wholesale prices.

According to the likelihood ratio test, performed in the basic model, the log-linear form was preferred to the linear form. However, all the minimization algorithms which are utilized by Hendry's program, for the estimation of general autoregressive schemes of the disturbances, failed to produce acceptable results when we used logarithms of the variables, because of different local minima. Therefore, the linear form was adopted for the imported commodity group under consideration.

Using seasonally adjusted data, the least squares estimates of the basic equation are as follows:

(1) $$M_{mt} = \underset{(12.838)}{39.258} - \underset{(0.091)}{0.426} \, (P^{m}_{mt}/P^{d}_{g}) + \underset{(0.034)}{0.940} \, IP$$

$n=263$, $RSS=43918.208$, $R^2=0.928$, $D.W.=1.561$, $X^2(15)=84.460$

where,

M_{mt} = Index of volume of imported machinery and transport equipment without ships

P^{m}_{mt} = Price index of imported machinery and transport equipment without ships

P_g^d - Domestic general index of wholesale prices

IP - Index of industrial production

The activity and price regression coefficients obtained from the above estimated equation are highly significant and they have the correct sign. The Durbin-Watson statistic is significant revealing a first order serial dependence in the residuals. The $X^2(15)$ Box-Pierce test-statistic has a highly significant value indicating that higher order autocorrelations are also present in the residuals. In fact, there are peaks in the residual correlogram at lags 1, 2, 3, 6, 7 and 11. So we reestimated the above equation allowing the error term to follow a simple ℓth order autoregressive process and experiments were made for values of ℓ from 1 to 13. As it was expected, according to the likelihood ratio or the asymptotically equivalent t-test, we obtained a significant autocorrelation parameter only for the values of ℓ corresponding to the above peaks in the residual correlogram. In all cases the estimated structural coefficients were very close to the ones of equation (1) and, except for ℓ = 2 and 7, the original equation, autoregressive error hypothesis, was rejected in favor of the unrestricted transformed equation. The above indicate that a general ℓth order autoregressive error process would be more appropriate, which finally was preferred, and in likelihood ratio tests the hypothesis that ℓ = 3 was accepted against other alternative values.

The simultaneous estimates of the regression coefficients and the autocorrelation parameters are:

$$(2) \qquad M_{mt} = \underset{(14.411)}{46.180} - \underset{(0.101)}{0.470}\,(P^{m}_{mt}/P^{d}_{g}) + \underset{(0.045)}{0.921}\,IP$$

$$u_t = \underset{(0.060)}{0.164}\,u_{t-1} + \underset{(0.061)}{0.058}\,u_{t-2} + \underset{(0.060)}{0.292}\,u_{t-3}$$

$$n = 263,\ \text{R.S.S.} = 37802.252,\ X^2(15) = 15.682$$

Turning now to a test of the non-linear restrictions imposed in estimation we find that the original specification, autoregressive error hypothesis, is rejected in favor of the unrestricted transformed equation ($X^2(6) = 13.044$). This equation, containing one, two and three months lagged values of all three variables and an independent error term, achieves a significantly lower residual sum of squares than the restricted form given above. So the preferred form, after dropping the non-significant variables, is

$$(3) \qquad M_{mt} = \underset{(12.064)}{28.079} + \underset{(0.057)}{0.168}\,M_{mt,t-1} + \underset{(0.056)}{0.274}\,M_{mt,t-3} -$$

$$- \underset{(0.087)}{0.283}\,(P^{m}_{mt}/P^{d}_{g}) + \underset{(0.073)}{0.522}\,IP$$

$$n=263,\ RSS= 37676.407,\ R^2=0.938,\ D.W.=1.971,\ X^2(15)=19.645$$

Following now the common factor analysis we begin with the general unrestricted dynamic model of the form

$$(4) \qquad a(L)M_{mt} = \beta_1(L)(P^{m}_{mt}/P^{d}_{g}) + \beta_2(L)IP + \text{constant}$$

where the scalar polynomials in the lag operator L are of degree 13. The least squares estimates of (4) gave an equation with 42 coefficients from which only the coefficients of M_{mt} lagged one and three months as well as the coefficients of IP lagged four months and (P^{m}_{mt}/P^{d}_{g}) came out significant and with the correct sign ($n = 263$, R.S.S. $= 30889.883$, $R^2 = 0.949$, D.W. $= 2.036$, $X^2(15) = 3.114$).

The table below gives the Wald criteria, as well as their successive differences, for a common factor polynomial of degree m and for m = 10, 11, 12, and 13.

Wald Criteria

m	Wald Criterion	Degrees of Freedom	Diffe-rences	Degrees of Freedom
10	1.514	20	-	-
11	1.627	22	0.113	2
12	10.871	24	9.244*	2
13	43.306*	26	32.435*	2

* Significant at the 2% level.

Taking first the Wald criteria, it seems fairly clear from the above table that the hypothesis that m = 13 should be rejected in favor of m = 12. This suggests that we require a structural equation form in which at least some of the variables of equation (1) have one extra lag. Reestimating the basic equation including all variables lagged one month, only the lagged dependent variable produced a significant coefficient and on estimation subject to an autoregressive error term, a simple 3rd order autoregressive scheme was preferred.

The estimated equation is:

$$(5) \qquad M_{mt} = \underset{(13.091)}{42.635} + \underset{(0.058)}{0.208}\, M_{mt,t-1} - \underset{(0.093)}{0.418}\, (P^{m}_{mt}/P^{d}_{g}) + \underset{(0.070)}{+\ 0.720}\, IP$$

$$u_t = \underset{(0.060)}{0.310}\, u_{t-3} + \varepsilon_t$$

$n = 263$, R.S.S. $= 37407.591$, $X^2(15) = 17.548$

which being accepted against the unrestricted transformed equation ($X^2(3) = 5.641$) is the preferred form according to the Wald criteria.

Taking now the Wald difference criteria, it appears from the above table that, on the basis of the difference between m = 12 and m = 13, and between m = 11 and m = 12, the hypothesis that m = 13 or 12 might be rejected in favor of m = 11. This indicates that we require a structural equation form in which at least some of the variables of equation (1) have two extra lags (or one value lagged two months). Estimating the basic equation including this time all variables lagged one and two months, only the lagged values of the dependent variable produced a significant coefficient and reestimating allowing the error term to be autoregressed, a simple 3rd order autoregressive form was preferred.

The maximum likelihood estimates are:

$$(6) \qquad M_{mt} = \underset{(13.042)}{42.195} + \underset{(0.060)}{0.188}\, M_{mt,t-1} + \underset{(0.059)}{0.088}\, M_{mt,t-2} -$$

$$- \underset{(0.093)}{0.406}\, (P^m_{mt}/P^d_g) + \underset{(0.083)}{0.654}\, IP$$

$$u_t = \underset{(0.061)}{0.299}\, u_{t-3} + \varepsilon_t$$

$n = 263$, R.S.S. $= 37091.911$, $X^2(15) = 14.477$

The test of the non-linear restrictions imposed in estimation, gave a value of 5.087 for the $X^2(4)$ test-statistic, which is not significant; hence, the original equation, autoregressive error specification is accepted against the unrestricted transformed equation and equation (6) is the preferred form according to the Wald difference criteria. Notice, however, that the coefficient of $M_{mt,t-2}$ came out insignificant and if we drop that variable, equation (5) is again the preferred form.

As can be seen from the above estimated equations (5) and (6), as were suggested by the Wald and difference Wald criteria respectively, the increase of the systematic dynamics, by one in the first case and by two in the second, did not affect the order of the error dynamics. In other words, in both equations a significant 3rd order error autoregression was found, and this probably reflects the omission of $M_{mt,t-3}$, from the above structural equations, which according to the likelihood ratio test, mentioned before in relation with specification (2), it should have been included. On the other hand, testing the significance of the lagged variables in the general unrestricted dynamic model (4), we found that the overall dynamics of the equation is reduced from 13 to 3 months (the F-test on the coefficients of the variables lagged four up to thirteen months gives a value of $F(30,210) = 1.149$ which is not significant). Applying the "common factor" analysis in the simplified version of (4) we obtained the following Wald criteria.

Wald Criteria

m	Wald Criterion	Degrees of Freedom	Diffe- rences	Degrees of Freedom
1	0.351	2	0.351	2
2	1.823	4	1.472	2
3	18.524*	6	16.701*	2

*Significant at the 1% level

Here again, it seems fairly clear that, on the basis of both the Wald and the difference Wald criteria, the hypothesis that m = 3 might be rejected in favor of m = 2 which leads us to accept equation (5) as the preferred form. But, if we reestimate the same structural equation allowing the error term to follow a 2nd order general autoregressive process, in order to retain the overall dynamics at three, we obtain the following equation:

$$(7) \qquad M_{mt} = \underset{(7.974)}{12.352} + \underset{(0.053)}{0.707}\, M_{mt,t-1} - \underset{(0.059)}{0.133}\, (P^{m}_{mt}/P^{d}_{g}) + \underset{(0.053)}{+ 0.281}\, IP$$

$$u_t = - \underset{(0.071)}{0.539}\, u_{t-1} - \underset{(0.067)}{0.322}\, u_{t-2} + \varepsilon_t$$

$$n = 263, \text{ R.S.S.} = 38504.903, \; X^2(15) = 17.091$$

which, while revealing a partial adjustment mechanism with a coefficient of adjustment smaller than the one of equation (5), in a test of the restrictions imposed in estimation, it is rejected in favor of the unrestricted transformed equation ($X^2(4) = 12.017$) which, after dropping the non-significant variables, gives the already estimated model (3). Finally, in terms of residual variance model (5) provides a slightly better fit giving s^2 144.991 instead of 146.033 of the model (3).

Turning now to a test of the seasonal adjustment of the data, we reestimated both equations (3) and (5) using unadjusted data. The F-tests constructed from the corresponding residual sums of squares gave the values of 2.023 and 2.166 respectively, for the $F(11,247)$ test-statistic which are significant. According to the likelihood ratio test, performed in both models (3) and (5), the log-linear form is preferred to the linear form. But since a complete empirical analysis using logarithms of the variables is not possible, as it was mentioned in the beginning of the section, we believe that the selected preferred form based on the levels of the variables should be retained.

The short-run income elasticities computed at the point of the sample means from models (3) and (5) are 0.639 and 0.882, whereas the corresponding long-run elasticities are 1.146 and 1.114 respectively. These long-run activity elasticities slightly above unity indicate that though the Greek industry has advanced considerably there is still a tendency for more mechanization and modernization. As far as imports of consumer durables are concerned, since, as already mentioned, industrial production has grown faster than national income, the income elasticity for these goods is still higher and that reflects their luxury character.

The short-run and long-run price elasticities computed at the point of the sample means are -0.498 and -0.893 for model (3), and -0.736 and -0.929 for model (5), respectively, all being highly significant. These below unity price elasticities, while being justifiable since the bulk of the imported commodities in this group is not produced at home, indicate that variations in relative prices exert a noticeable influence on the import demand for machinery and transport equipment.

2.7. Import Demand Equations for All Goods Excluding Ships (SITC: Sections 0-9)

Thus far, we have estimated import demand equations for the major groups of imported commodities which constitute about 96 percent in the total value of imported goods. In this section we try to estimate a demand equation for the imported commodities taken together.

According to the discussion in the previous sections, the volume of total imports is assumed to depend on the level of industrial production, and the price of imports, adjusted for import duties and taxes, relative to domestic wholesale prices.

For the basic model, Sargan's maximum likelihood ratio gave a value of 1.426 indicating that the log-linear form is preferred to the linear form. Using seasonally adjusted data, the least squares estimates of the basic equation are as follows:

$$(1) \qquad \log(M_T) = \underset{(0.559)}{2.331} - \underset{(0.101)}{0.520} \log(P^m_T/P^d_g) + \underset{(0.023)}{0.999} \log(IP)$$

$n=263$, $RSS=3.827$, $R^2=0.962$, $D.W.=1.202$, $X^2(15)=233.371$

where,

M_T = Index of volume of total imports

P_T^m = Price index of total imports

P_g^d = Domestic general index of wholesale prices

IP = Index of industrial production .

It appears from the above estimated equation that the regression coefficients have the expected signs and they are statistically significant. Both the Durbin-Watson and the $X^2(15)$ Box-Pierce test-statistics are highly significant indicating that first order as well as higher order autocorrelations are present in the residuals. In fact there are autocorrelations in the residual correlogram at all lags, taking values in the range 0.143 to 0.398. The above indicate that a general autoregressive error scheme should be considered. So, we reestimated equation (1) allowing the error term to follow a general ℓth order autoregression which finally was preferred to a simple autoregressive form, and in likelihood ratio tests the hypothesis that $\ell = 9$ was accepted against other values of ℓ.

The maximum likelihood estimates are:

$$(2) \qquad \log(M_T) = \underset{(0.867)}{5.701} - \underset{(0.154)}{1.068} \log(P_T^m/P_g^d) + \underset{(0.077)}{0.820} \log(IP)$$

$$u_t = \underset{(0.063)}{0.276}\, u_{t-1} + \underset{(0.065)}{0.141}\, u_{t-2} + \underset{(0.065)}{0.156}\, u_{t-3} +$$

$$+ \underset{(0.066)}{0.001}\, u_{t-4} - \underset{(0.065)}{0.032}\, u_{t-5} + \underset{(0.066)}{0.065}\, u_{t-6} -$$

$$- \underset{(0.065)}{0.017}\, u_{t-7} + \underset{(0.065)}{0.140}\, u_{t-8} + \underset{(0.063)}{0.127}\, u_{t-9}$$

$$n = 263,\ R.S.S. = 2.734,\ X^2(15) = 6.180\ .$$

We notice that the income elasticity decreased in value whereas the price elasticity increased in absolute terms. The test of the non-linear restrictions imposed by the auto-regressive error specification (2) in the corresponding unrestricted transformed equation, which contains one up to nine months lagged values of all three variables, gave a value of 26.599 for the $X^2(18)$ test statistic which is not significant; hence, the original equation, autoregressive error specification, is preferred.

Following now the "common factor" analysis, the general unrestricted dynamic model has the form

$$(3) \qquad a(L)\log(M_T) = \beta_1(L)\log(P_T^m/P_g^d) + \beta_2(L)\log(IP) + \\ + \text{constant}$$

where the scalar polynomials in the lag operator L are of order 13. The least squares estimates of (3) gave an equation with 42 coefficients, from which only the coefficients of the dependent variable lagged one, three and eight months as well as the coefficient of $\log(P_T^m/P_g^d)_t$ and $\log(IP)_t$ came out significant ($n = 263$, R.S.S. = 2.342, $R^2 = 0.977$, D.W. = 1.938, $X^2(15) = 2.859$). Testing for a common factor polynomial of degree m, and for m = 11, 12 and 13, we obtained the following Wald and difference Wald criteria.

Wald Criteria

m	Wald Criterion	Degrees of Freedom	Diffe-rences	Degrees of Freedom
11	2.334	22	-	-
12	4.214	24	1.880	2
13	15.508	26	11.294*	2

*Significant at the 1% level

Taking first the Wald criteria, none of them takes a significant value and therefore the hypothesis that m = 13 is accepted. This suggests that a structural equation form without any lags on the variables and a general autoregressive error specification should be considered, and the already estimated model (2) is the preferred form.

Taking now the Wald difference criteria, it appears from the above table that, on the basis of the difference between m = 12 and m = 13, the hypothesis that m = 13 might be rejected in favor of m = 12. This indicates that we require a structural equation form in which at least some of the variables of equation (1) have one extra lag. Reestimating the basic equation including all variables lagged one month, only the lagged dependent variable produced a significant coefficient and on estimation subject to an autoregressive error term, a general 9th order autoregressive scheme was preferred.

The estimated equation is:

$$(4) \qquad \log(M_T) = \underset{(0.951)}{6.094} - \underset{(0.113)}{0.117} \log(M_T)_{t-1} - \underset{(0.161)}{1.120} \log(P^m_T/P^d_g) + \underset{(0.123)}{0.900} \log(IP)$$

$$u_t = \underset{(0.131)}{0.391} u_{t-1} + \underset{(0.093)}{0.088} u_{t-2} + \underset{(0.068)}{0.142} u_{t-3} - \underset{(0.069)}{0.009} u_{t-4} - \underset{(0.067)}{0.026} u_{t-5} + \underset{(0.068)}{0.056} u_{t-6} - \underset{(0.068)}{0.021} u_{t-7} + \underset{(0.068)}{0.143} u_{t-8} + \underset{(0.067)}{0.113} u_{t-9}$$

$$n = 263, \text{ R.S.S.} = 2.720, \ X^2(15) = 6.392$$

which being accepted against the unrestricted transformed equation ($X^2(18) = 25.402$) is the preferred form according to the Wald difference criteria. Notice, however, that the coefficient of $\log(M_T)_{t-1}$ came out insignificant and if we drop that variable, equation (2) is again the preferred form.

The previously preferred autoregressive error specification (2) shows that the initially extracted set of 13 common roots (according to the Wald criteria, or 12 common roots according to the difference Wald criteria) contains four zeros which in the general unrestricted dynamic model (3), shorten the lag length by four for every variable. This was confirmed in a F-test on the coefficients of the variables lagged 10 up to 13 months, which gave a value of 0.967 for the $F(12,210)$ test-statistic which is not significant. Applying the "common factor" analysis in the simplified version of (3) we obtained the following Wald criteria.

Wald Criteria

m	Wald Criterion	Degrees of Freedom	Differences	Degrees of Freedom
7	1.421	14	-	-
8	4.063	16	2.642	2
9	14.495	18	10.432*	2

*Sifnificant at the 1% level

The results are similar to the ones obtained before from the application of the common factor analysis in the general unrestricted dynamic model (3). In particular, the Wald criteria suggest that a structural equation form without any lags on the variables and an autoregressive error specification of general 9th order form, i.e. model (2), should be considered as the data generation process. But, on the basis of the difference between m = 8 and m = 9, the hypothesis that m = 9 might be rejected in favor of m = 8 indicating that first order systematic dynamics should be introduced into the structural equation, which finally lead to the specification (4). In this case the overall dynamics in the equation is increased from 9th order to 10th order. If we reestimate the same structural equation allowing the error term to follow a 8th order general autoregressive process, in order to retain the overall dynamics at nine, we obtain the following equation:

$$(5) \qquad \log(M_T) = \underset{(0.905)}{5.875} - \underset{(0.103)}{0.091} \log(M_T)_{t-1} -$$

$$- \underset{(0.151)}{1.125} \log(P^m_T/P^d_g) + \underset{(0.112)}{0.928} \log(IP)$$

$$u_t = \underset{(0.119)}{0.387}\, u_{t-1} + \underset{(0.086)}{0.101}\, u_{t-2} + \underset{(0.068)}{0.147}\, u_{t-3} -$$

$$- \underset{(0.068)}{0.015}\, u_{t-4} - \underset{(0.067)}{0.030}\, u_{t-5} + \underset{(0.067)}{0.069}\, u_{t-6} -$$

$$- \underset{(0.067)}{0.010}\, u_{t-7} + \underset{(0.063)}{0.182}\, u_{t-8}$$

$$n = 263, \text{ R.S.S.} = 2.762, \ X^2(15) = 7.692$$

which, while giving similar values for the structural coefficients and autocorrelation parameters to equation (4), in a test of the restrictions imposed in estimation, it is rejected in favor of the unrestricted transformed equation ($X^2(16) = 29.128$).

Turning now to a test of the seasonal adjustment of the data, we reestimated equation (2) using unadjusted data. The F-test constructed from the corresponding residual sums of squares gave a value of 9.519 for the $F(11,240)$ test statistic which is significant. Finally, according to the likelihood ratio test, performed in model (2), the log-linear form is preferred to the linear form.

The activity elasticity of demand for the total imports, estimated by equation (2), is 0.82 whereas the estimated price elasticity is -1.068, both being highly significant. These estimates, derived from an aggregate demand function, are likely to be subject to aggregation bias. In order to determine whether a significant difference emerges on calculating import demand elasticities from a series of disaggregated demand functions rather than from an aggregated demand function we have followed the method used by Barker (1970) after transforming it to fit our data.

To illustrate the sources of bias in log-linear aggregation, let us consider the following import demand function for the ith group of commodities

$$(6) \qquad M_i = a_{0i} \, Y_i^{a_{1i}} \, P_i^{a_{2i}}$$

where, M_i is imports in category i, Y_i is total demand in category i (or the ith component of total income) and P_i is the relative price variable.

The demand elasticity is defined as

$$(7) \qquad \frac{\partial M_i}{\partial Y_i} \cdot \frac{Y_i}{M_i} = a_{1i}$$

and the relative price elasticity as

$$(8) \qquad \frac{\partial M_i}{\partial P_i} \cdot \frac{P_i}{M_i} = a_{2i}$$

From (7) and (8), the total variation in M_i is

$$\frac{dM_i}{M_i} = \frac{dY_i}{Y_i} a_{1i} + \frac{dP_i}{P_i} a_{2i}$$

Now the relative variation in total imports M, given $M = \sum_i M_i$, is

$$(9) \qquad \frac{dM}{M} = \sum_i \frac{dM_i}{M} = \sum_i \frac{dM_i}{M_i} \frac{M_i}{M} =$$

$$= \sum_i \frac{dY_i}{Y_i} a_{1i} \frac{M_i}{M} + \sum_i \frac{dP_i}{P_i} a_{2i} \frac{M_i}{M}$$

The overall demand (or income) elasticity is found, using the first term of equation (9), as

$$a_1 = \frac{\partial M}{\partial Y} \cdot \frac{Y}{M} = \sum_i a_{1i} \left(\frac{M_i}{M}\right) \cdot \left(\frac{dY_i/Y_i}{dY/Y}\right)$$

(where Y is total demand or total income). That is, the overall demand elasticity depends on the individual ones (a_{1i}), the shares of the individual imports in total imports (M_i/M) and the relative variation in individual demand divided by that of total demand $\left(\frac{dY_i/Y_i}{dY/Y}\right)$. This last term is the "distribution" elasticity of demand and is obtained by regressing the percentage change in Y_i, (dY_i/Y_i), on the percentage change in Y, (dY/Y).

Similarly, the overall relative price elasticity is found, using the second term of equation (9), as

$$a_2 = \frac{\partial M}{\partial P} \cdot \frac{P}{M} = \sum_i a_{2i} \left(\frac{M_i}{M}\right) \left(\frac{dP_i/P_i}{dP/P}\right)$$

where, $\left(\frac{dP_i/P_i}{dP/P}\right)$ is similarly the "distribution" price elasticity of demand and is obtained by regressing the percentage change in the ratio of import to domestic price indexes of the ith commodity group (dP_i/P_i) on the percentage change in the relative price of total imports (dP/P).

Let us consider now a slightly different import demand function which we have used in our empirical analysis, namely

$$Q_i = \beta_{0i} \, I^{\beta_{1i}} \, P_i^{\beta_{2i}}$$

where, Q_i is an index of volume of imports in category i and I is an activity variable (index of industrial production) which has been used as income-proxy in all our import demand equations. Let Q denote the index of volume of total imports, which is a weighted index of the Q_i so that $Q = \sum_i W_i Q_i$, where W_i are the weights. Then, following the same analysis as before, the relative variation in Q is

$$(10) \qquad \frac{dQ}{Q} = \sum_i W_i \frac{Q_i}{Q} \cdot \frac{dI}{I} \beta_{1i} + \sum_i W_i \frac{Q_i}{Q} \cdot \frac{dP_i}{P_i} \beta_{2i}$$

Using the first term of equation (10), the overall income elasticity is found as

$$(11) \qquad \beta_1 = \frac{\partial Q}{\partial I} \cdot \frac{I}{Q} = \sum_i W_i \frac{Q_i}{Q} \beta_{1i} \frac{dI}{I} \cdot \frac{I}{dI} =$$

$$= \sum_i W_i \frac{Q_i}{Q} \beta_{1i}$$

Similarly, the overall relative price elasticity is found, using the second term of equation (10), as

$$(12) \qquad \beta_2 = \frac{\partial Q}{\partial P} \cdot \frac{P}{Q} = \sum_i W_i \frac{Q_i}{Q} \beta_{2i} \left(\frac{dP_i/P_i}{dP/P} \right)$$

where, $\frac{dP_i/P_i}{dP/P}$ is, as before, the "distribution" price elasticity of demand.

Relations (11) and (12), from which the aggregate income and relative price elasticities are estimated, include the products $W_i \frac{Q_i}{Q}$. As already mentioned in section 1, our volume indices are of the Laspeyres type. Let $m_{i,t}$ and m_t denote the value of imports of the ith group of commodities, and the value of total imports respectively, at period t, both valued at base period (0) prices (in our data 0 = 1970). Then

$$Q_t = \frac{m_t}{m_0} = \frac{\sum_i m_{i,t}}{m_0} = \sum_i \frac{m_{i,0}}{m_0} \cdot \frac{m_{i,t}}{m_{i,0}} =$$

$$= \sum_i W_i \; Q_{i,t}$$

and therefore

$$W_i \frac{Q_i}{Q} = \frac{m_{i,0}}{m_0} \cdot \frac{m_{i,t}}{m_{i,0}} \cdot \frac{m_0}{m_t} = \frac{m_{i,t}}{m_t}$$

So, equations (11) and (12) become

(13) $$\beta_1 = \sum_i \frac{m_{i,t}}{m_t} \beta_{1i}$$

and

(14) $$\beta_2 = \sum_i \frac{m_{i,t}}{m_t} \beta_{2i} \left(\frac{dP_i/P_i}{dP/P} \right)$$

It appears then, that the overall activity elasticity depends only on the individual ones (β_{1i}) and the shares of the individual imports in total imports. The distribution income elasticity of demand does not appear since the same activity variable enters all import demand equations. The overall relative price elasticity is estimated from equation (14) in the same way as previously described. Notice that the shares of the individual imports in total imports ($m_{i,t}/m_t$) at period t, result from the corresponding values of individual and total imports at that period, but valued at base period (1970) prices.

The results obtained from using equations (13) and (14) to estimate the short-run aggregate import demand elasticities from the disaggregated functions estimated before, are presented in the table below. Since the preferred model for

TABLE 1

INCOME AND PRICE ELASTICITIES FOR SIX CATEGORIES OF IMPORTS AND TOTAL IMPORTS

Import Category	Short - run Income Elasticity	Short - run Price Elasticity	Shares in Total Imports		Estimated Relative Price Distribution Elasticity
			Average 1954-76	1976	
Food	0.460	-0.805	0.133	0.099	0.763 (0.126)
Crude Materials	0.939	-0.264	0.104	0.099	0.244 (0.083)
Fuels	0.690	-0.325	0.094	0.105	1.113 (0.173)
Chemicals	1.000	-1.269	0.098	0.107	0.619 (0.134)
Manufactures	0.793	-0.937	0.223	0.202	0.342 (0.074)
Machinery and Transport Equipment	1.128	-0.828	0.309	0.336	0.793 (0.091)
Total direct estimate	0.820	-1.068			
Total estimate derived using					
(i) Average shares	0.847	-0.474			
(ii) 1976 shares	0.857	-0.475			

over-all imports is of the static form (model (2)), the disaggregated import demand functions we consider are of the same form. In particular, for the categories food, crude materials, chemicals and manufactures, the elasticities result from the corresponding preferred original equations, autoregressive error specifications, already mentioned in the relevant sections. The import demand elasticities for fuels are obtained from the log-linear version of the preferred linear form.

The elasticities of import demand for machinery and transport equipment are obtained from the preferred original equation, autoregressive error specification which is in linear form since, as already mentioned, convergence difficulties made impossiple the estimation of general autoregressive error forms when we used logarithms of the variables. Finally, aggregate elasticities are derived, using the average and the 1976 shares of each import category in the total.

From the above table it can be seen that with respect to the activity variable, there is no important difference between the direct estimate of the aggregate import demand elasticity and the estimate derived from the disaggregated functions.

But with respect to the price variable, a considerable difference emerges on computing the aggregate import demand elasticity from the disaggregated functions rather than by direct estimation. Thus one finds that the price elasticity of demand for over-all imports now appears to be in the region of -0.5 rather than -1.1. This is due to the fact that those imports with relatively higher price elasticities, though with larger shares in total imports, have low import price distribution elasticities. If we ignore the price distribution elasticities, the weighted aggregate price elasticity, using the average and the 1976 shares, is -0.754 and -0.743 respectively.

So, in spite of Orcutt's (1950) arguments against aggregation, it appears that the price elasticity estimated by the aggregate import demand equation is biased upwards. Though it is not a general conclusion about aggregation bias, this result demonstrates that there are cases where aggregation leads to an overestimate of the price elasticity and also reveals the presence of bias in estimations using aggregate data.

3. Summary of Findings

A summary of the estimated elasticities of the import demand equations for the major groups of imported commodities and total imports is presented in the table below. Long-run income and price elasticities are presented for those groups of commodities for which dynamic import demand functions were estimated.

Since industrial production has grown faster than national income, the import demand elasticity with respect to national income is expected to be somewhat higher than the import demand elasticity with respect to industrial production. But monthly data are not available for national income. Therefore, for purposes of comparison, we have tried to approximate the national income elasticity indirectly by using

TABLE 2

IMPORT DEMAND ELASTICITIES FOR GOODS
(Monthly Data 1954 - 1976)

Imports of:	Industrial Production Elasticity		Relative Price Elasticity		National Income Elasticity	
	Short-run	Long-run	Short-run	Long-run	Short-run	Long-run
Food	0.460	0.467	-0.805	-0.775	0.619	0.628
Crude Materials	0.939	0.964	-0.264[a]	-0.316[a]	1.263	1.297
Fuels						
1) Linear form	0.906		-0.264		1.219	
2) Log-linear form	0.690	0.615	-0.325	-0.541	0.928	0.827
Chemicals	0.673[b]	0.604[b]	-1.277	-1.876		
	1.000				1.345	
Manufactures	0.793	0.799	-0.937	-1.078	1.067	1.075
Machinery and Transport Equipment	1.128	1.114	-0.828	-0.929	1.517	1.498
All Goods	0.820		-1.068		1.103	

[a]Not significant at the 5 percent level

[b]Chemical production elasticity

the "function of a function" rule (see Paraskevopoulos (1970), p. 108). That is, to obtain the national income elasticity, the industrial production elasticity was multiplied by 1.345[1] - the estimated elasticity for industrial production with respect to national income.

The major findings of the preceding empirical analysis may be summarized as follows:

(i) Imports of food are treated as essentials by the Greek consumer. The increasing real income and the rise in the standard of living are the main reasons that imported food items, although of higher quality, are not treated as luxuries by Greek consumers. The per capita consumption of these food items (meat, fish, dairy products etc.) in Greece is quite close to that experienced by high income countries.

(ii) The volume of imported raw materials and fuels has grown as fast as the level of industrial production. Greece is a very small country and relatively poor in resource endowment. Hence, imports of raw materials and fuels are indispensable for the industrialization of the country and any attempt to interrupt the flow of these imports would undoubtedly hold back the efforts of economic development.

(iii) The volume of imported raw chemical materials has grown slower than the level of chemical production. Domestic chemical industry, which includes among other things the production of raw chemical materials, has advanced

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[1] $\log(IP) = -2.854 + 1.345 \log(NI)$, $R^2 = 0.994$

rapidly resulting in a gradual shift of the demand from imported chemical raw materials to domestically produced ones. But as far as imported chemical manufactures are concerned, the income elasticity appears to be above unity.

(iv) The industrial production elasticity for imported capital goods is slightly above unity indicating that though the Greek industry has advanced considerably there is still a tendency for more mechanization and modernization. On the other hand the income elasticity for imported consumer durables is clearly above unity, reflecting their luxury character, while for imported manufactures the income elasticity appears to be in the neighborhood of one.

(v) Finally, taking the relative price elasticities at their face value, their sizes on the whole, save the relative price elasticity for raw materials, indicate that movements in relative prices are quite important in determining changes in the volume of imported commodities.

CHAPTER VI

THE ESTIMATED EXPORT DEMAND EQUATIONS OF GREECE

In the preceding chapter, the import demand functions of Greece have been estimated. In the present chapter, we shall try to estimate the foreign demand equations for the Greek exports of goods.

1. Statistical Data

The data for exports used in this study are time series observations for the period 1954 - 1976. This selection was largely dictated by the presentation and reliability of the basic data. As already mentioned in chapter V, we have decided to use monthly data which are available only for the period 1954 onwards.

Exports, for the purpose of statistical analysis, have been broken down into five major groups of commodities: food, beverages and tobacco, crude materials, chemicals, and manufactures. This particular classification coincides with the Standard International Trade Classification adopted by the United Nations; it was dictated by the presentation of the basic data, i.e. foreign trade indices, which are published by the Statistical Service of Greece at the one digit SITC level of aggregation (the above groups are the corresponding SITC categories 0, 1, 2, 5 and 6). The omitted SITC categories (3, 4, 7, 8 and 9) are not examined because being unimportant for Greek export trade, they are not listed for the period before 1972. They are, however, included in total exports.

The basic statistical data of Greek exports were taken primarily from the Monthly Bulletin of External Trade Statistics published by the National Statistical Service of Greece. Exports are valued at free on board (f.o.b.) basis, and are recorded at the clearance point of the Customs authorities. The trade statistics give volume indices of the Laspeyres type as well as price indices of the Paasche type for the total exports and the major groups of commodities mentioned above. The price indices are unit values obtained as the ratio of the current value of exports to their value at constant prices. Since the indices are given in three separate series based on 1954, 1961 and 1970 prices respectively, continuous series for the sample period were obtained by splicing together the three separate series.

The basic hypothesis underlying the export demand equations is in a way analogous to that used for the import demand equations. Variations in the volume of exports are mainly assumed to depend on changes in income or activity of the foreign countries and movements in export prices relative to the prices of related commodities in foreign markets.

The exports of Greece are concentrated in a few developed countries which are rather homogeneous as far as demand conditions are concerned. The O.E.C.D. countries, and particularly West Germany, United States, Italy, France, United Kingdom, Yugoslavia and Netherlands, are the primary customers of Greece taking more than 70 percent of the country's exports, while the remainder is roughly divided between the Eastern European countries and the rest of the world. The above mentioned seven O.E.C.D. countries taken together absorb more than 60 percent of Greece's exports and their average shares in the value of the total exported goods during the period 1954 - 1976 were as follows; West Germany: 20.8%, U.S.A.: 10.7%, Italy: 8.5%, France: 7.2%, U.K.: 7.1%, Yugoslavia: 4.3%, and Netherlands: 3.8% .

It was decided, therefore, that the income or activity variable to be used should be that of the total of O.E.C.D. countries. But for the prices of related commodities in foreign markets, we consider only the ones of these seven countries.

The principal sources of the above data are the Main Economic Indicators - Historical Statistics, 1955 - 1971, and 1960 - 1975, respectively, as well as various issues of the Main Economic Indicators, all published by O.E.C.D. Since monthly figures for the income of the O.E.C.D. countries are not published, the index of industrial production of the O.E.C.D. countries or its appropriate component has been selected as an income-proxy.

The export prices were deflated by the index of domestic prices of the importing countries, constructed from the wholesale prices of related commodities in these seven countries, weighted according to their (annual) shares in Greece's total exports. It should be mentioned that the indices of domestic prices published by and from the O.E.C.D. countries are not homogeneous, and so for each country we have selected the price series which is most related to the particular group of exported commodities.

Finally, apart from price and activity variables a time trend was also included in the export demand equations to take account of any systematic changes in other influences which have not been introduced explicitly into the equations as for example changes in tastes, technology and the like. However this time variable failed to produce significant results because it was highly intercorrelated with the activity variables.

2. The estimated Export Demand Equations of Greece for Major Groups of Commodities

2.1. Export Demand Equations for Food (SITC: Section 0)

Exports of food amounted to about 22 to 26 percent in the value of total exports of the country for the period 1954 - 1976. This export group mainly includes fruit and nuts, fresh or dried (but mainly dried), vegetables, fresh and preserved, and preserved fruit, which account for more than 80 percent in the exports of food. Though the share of food in the total exports of goods remained almost unchanged over the period 1954 - 76, the composition of exported food has been changed considerably. In particular, the share of fruit and nuts (fresh or dried) in the total exports of food decreased from 84.2% in 1954 to 40.2% in 1976. On the contrary exports of more dynamic food items such as vegetables (fresh and preserved) and preserved fruit, which accounted for 2% and 9.9% in the exports of food in 1954 respectively, increased to 25.8% and 14.9% in 1976.

In 1966, for the first time, Greece exported wheat which accounted for 22.5% of the total exports of food and declined ever since to reach 5.4% in 1976. The remainder of the group includes mainly, cheese, fish fresh and animal feeds.

Since the majority of exported food items are not further processed in production, that is, they are objects of final consumption, the total index of industrial production of the O.E.C.D. countries, which is more closely related to the real income than any other index of industrial activity, has been selected as the income variable. To obtain the relative price variable the price index of exported food was deflated by the weighted index of domestic prices of related commodities of the seven previously mentioned importing countries. The particular price series used are the following:

U.S.A.	:	Price index of food (wholesale prices)
France	:	Price index of agricultural goods (producer prices)
Germany	:	Price index of food, beverages and tobacco (producer prices)
Italy	:	Price index of food (wholesale prices)
Netherlands	:	Price index of food (wholesale prices)
U.K.	:	Price index of food (wholesale prices)
Yugoslavia	:	Price index of food (producer prices)

For the basic model, Sargan's maximum likelihood ratio gave a value of 1.47 indicating that the log-linear form is preferred to the linear form. The least-squares estimates of the basic equation obtained from seasonally adjusted data, are as follows, with standard errors in parentheses:

$$(1) \qquad \log(X_f) = \underset{(0.662)}{-\,3.058} - \underset{(0.125)}{0.130} \log\left(\frac{P_f^x}{P_f^{OECD7}}\right) + \underset{(0.061)}{1.808} \log(IP_{OECD})$$

$n=263$, $RSS=26.062$, $R^2=0.775$, $D.W.=0.877$, $X^2(15)=211.450$

where,

X_f = Index of volume of exported food

P_f^x = Price index of exported food

P_f^{OECD7} = Weighted index of domestic prices of food of seven importing OECD countries

IP_{OECD} = Index of industrial production of OECD

It appears from the estimated equation that the income and the price coefficients have the expected signs, but the price coefficient is insignificant. Both Durbin-Watson and the $X^2(15)$ Box-Pierce test-statistics are highly significant indicating that first order as well as higher order autocorrelations are present in the residuals. In fact there are peaks in the residual correlogram at lags 1, 2, 5, 11, 12 and 13. Reestimating equation (1) subject to a simple ℓth order autoregressive error, we obtained, as expected, a significant autocorrelation parameter for the values of ℓ corresponding to the above peaks in the residual correlogram. In all cases the estimated activity elasticity was very close to the one of equation (1), whereas for $\ell = 11$, 12 and 13, the price elasticity increased in absolute value and became significant. Also, apart from $\ell = 1$, 2 and 5, in all other cases the original equation, autoregressive error specification was rejected in favor of the unrestricted transformed equation, whereas in all the aforementioned equations the $X^2(15)$ Box-Pierce test-statistic for a random correlogram gave a significant value.

The above indicate that a general autoregressive error scheme should be more appropriate. So we reestimated equation (1) allowing this time the error term to follow a general ℓth order autoregression which finally was preferred and in likelihood ratio tests the hypothesis that $\ell = 13$ was accepeted against other alternative values.

The maximum likelihood estimates are:

$$(2) \qquad \log(X_f) = \underset{(1.343)}{-0.933} - \underset{(0.173)}{0.532} \log\left(\frac{P_f^x}{P_f^{OECD7}}\right) + \underset{(0.256)}{1.763} \log(IP_{OECD})$$

$$\begin{aligned} u_t = {} & \underset{(0.063)}{0.587}\, u_{t-1} - \underset{(0.070)}{0.106}\, u_{t-2} - \underset{(0.069)}{0.003}\, u_{t-3} + \\ & + \underset{(0.069)}{0.068}\, u_{t-4} + \underset{(0.069)}{0.056}\, u_{t-5} - \underset{(0.069)}{0.045}\, u_{t-6} + \\ & + \underset{(0.069)}{0.028}\, u_{t-7} - \underset{(0.069)}{0.042}\, u_{t-8} - \underset{(0.069)}{0.010}\, u_{t-9} - \\ & - \underset{(0.069)}{0.065}\, u_{t-10} + \underset{(0.068)}{0.226}\, u_{t-11} + \underset{(0.069)}{0.316}\, u_{t-12} - \\ & - \underset{(0.063)}{0.172}\, u_{t-13} \end{aligned}$$

$$n = 263, \ \text{R.S.S.} = 12.745, \ X^2(15) = 7.073$$

The highly significant values of the first and twelve order autoregressive parameters ρ_1 and ρ_{12} indicate that the factorised error specification $(1-\rho_1 L)(1-\rho_{12}L^{12})u_t = \varepsilon_t$ would be relevant (or the unrestricted error specification $u_t = \rho_1 u_{t-1} + \rho_{12} u_{t-12} + \rho_{13} u_{t-13} + \varepsilon_t$ since ρ_{13} is also significant although at the same time being approximately equal to $-\hat{\rho}_1 \times \hat{\rho}_{12}$ implies that the factorised form would be more likely to occur). However, all the minimization algorithms which are utilized by RALS, failed to produce acceptable results because of different local minima and so the above error specifications were not considered.

Turning now to a test of the non-linear restrictions imposed in estimation, we find that the original specification, autoregressive error hypothesis (2), is rejected in favor of the unrestricted transformed equation ($X^2(26) = 48.196$). This equation, containing one up to thirteen months lagged values of all three variables and an independent error term, achieves a substantially lower residual sum of squares than the restricted form given above, and after dropping the non-significant variables we obtain

$$(3) \qquad \log(X_f) = \underset{(0.542)}{0.264} + \underset{(0.050)}{0.462} \log(X_f)_{t-1} + \underset{(0.056)}{0.209} \log(X_f)_{t-11} +$$
$$+ \underset{(0.060)}{0.251} \log(X_f)_{t-12} - \underset{(0.093)}{0.128} \log(P_f^X / P_f^{OECD7}) +$$
$$+ \underset{(0.123)}{0.165} \log(IP_{OECD})$$

$n=263$, $RSS=13.943$, $R^2=0.880$, $D.W.=1.713$, $X^2(15)=30.378$

Though the Durbin-Watson test-statistic is in the incoclusive range there is a peak at lag 1 in the residual correlogram, which causes the $X^2(15)$ Box-Pierce test-statistic to take a significant value (the Durbin-Watson test-statistic is probably biased upwards because of the presence of lagged values of the dependent variable among the regressors). Reestimating equation (3) subject to a first order autoregressive error term we obtain the following equation:

$$\text{(4)} \qquad \log(X_f) = \underset{(0.848)}{0.072} + \underset{(0.093)}{0.116} \log(X_f)_{t-1} + \underset{(0.055)}{0.179} \log(X_f)_{t-11} + \underset{(0.060)}{0.424} \log(X_f)_{t-12} - \underset{(0.145)}{0.249} \log(P_f^x/P_f^{OECD7}) + \underset{(0.196)}{0.537} \log(IP_{OECD})$$

$$u_t = \underset{(0.095)}{0.469}\, u_{t-1} + \varepsilon_t$$

$$n = 263,\ \text{R.S.S.} = 13.016,\ X^2(15) = 12.550$$

which is accepted against the unrestricted transformed equation ($X^2(3) = 2.381$). We notice that the coefficient of $\log(X_f)_{t-1}$ is insignificant and if we drop that variable we obtain

$$\text{(5)} \qquad \log(X_f) = \underset{(0.911)}{-0.109} + \underset{(0.054)}{0.168} \log(X_f)_{t-11} + \underset{(0.053)}{0.459} \log(X_f)_{t-12} - \underset{(0.157)}{0.282} \log(P_f^x/P_f^{OECD7}) + \underset{(0.157)}{0.704} \log(IP_{OECD})$$

$$u_t = \underset{(0.053)}{0.548}\, u_{t-1} + \varepsilon_t$$

$$n = 263,\ \text{R.S.S.} = 13.142,\ X^2(15) = 14.674$$

which is also accepted against the unrestricted transformed equation ($X^2(3) = 0.993$). So, according to the likelihood ratio tests, equation (5) is the preferred form.

Considering now the "common factor" analysis the estimated general unrestricted dynamic model, has the form

$$\text{(6)} \qquad a(L)\log(X_f) = \beta_1(L)\log(P_f^x/P_f^{OECD7}) + \beta_2(L)\log(IP_{OECD}) + \text{constant}$$

where the scalar polynomials in the lag operator L are of order 13 (n = 263, R.S.S. = 10.611, R^2 = 0.908, D.W. = 2.007, $X^2(15)$ = 5.799).

The Wald test for a common factor polynomial of degree 13, gives a value of 14.975 for the $X^2(26)$ test statistic which is not significant. This suggests that a structural equation form without any lags on the variables and an autoregressive error specification of general 13th order form, i.e. model (2), should be considered as the data generation process.

The Wald test for a common factor polynomial of degree 12 gives the (non-significant) value of 5.585 for the $X^2(24)$ test-statistic. Thus, the Wald difference criterion is equal to 9.390, and exceeds 5.99 which is the 5% confidence limit for a $X^2(2)$, suggesting that we require a structural equation form in which some of the variables of equation (1) have one extra lag. Reestimating the basic equation including this time all the variables lagged one month, only the lagged dependent variable produced a significant coefficient and on estimation subject to an autoregressive error term, a simple 12th order autoregressive form was preferred.

The estimated equation is:

$$(7) \qquad \log(X_f) = \underset{(0.644)}{-\,0.384} + \underset{(0.052)}{0.523}\log(X_f)_{t-1} - \underset{(0.114)}{0.304}\log(P_f^x/P_f^{OECD7}) + \underset{(0.129)}{0.881}\log(IP_{OECD})$$

$$u_t = \underset{(0.055)}{0.498}\, u_{t-12} + \varepsilon_t$$

n = 263, R.S.S. = 13.635, $X^2(15)$ = 22.680

which according to the Wald difference criterion is the preferred form. However, in a likelihood ratio test, specification (7) is rejected in favor of the unrestricted transformed equation ($X^2(3) = 13.611$).

The above empirical analysis indicates that the overall dynamics in the export demand equation for food is of 13th order. The model selection procedure which starts from the static form of the basic equation leads to a model with 12th order systematic dynamics and a first order error dynamics (model (5)). On the contrary, starting from the general unrestricted dynamic model, the common factor analysis based on the Wald criteria suggests that the overall dynamics in the equation is expressed only through error dynamics (model (2)). But, the same analysis based on the Wald difference criteria, suggests that the overall dynamics is factorized into a first order systematic dynamics and a 12th order error dynamics (model (7)). In terms of residual variance model (5) provides a slightly better fit giving s^2 0.0511 instead of 0.0516 and 0.0528 of models (2) and (7) respectively.

Turning now to a test of the seasonal adjustment of the data, we reestimated all equations (2), (5) and (7) using unadjusted data. The F-tests constructed from the corresponding residual sums of squares gave the values of 3.009, 3.145 and 2.752 for the $F(11,236)$, $F(11,246)$ and $F(11,247)$ test statistics respectively, which are significant. Finally, according to the likelihood ratio test performed in all the above models, the log-linear form is preferred to the linear form.

The short-run income elasticities estimated from models (2), (5) and (7) are 1.763, 0.704 and 0.881 respectively, whereas the long-run income elasticities obtained from models (5) and (7) are 1.888 and 1.848 respectively. It appears then that with respect to the activity variable of the O.E.C.D. countries, the Greek exports of food are income elastic.

With respect to the price variable, the estimated short-run elasticities from models (2), (5) and (7) are -0.532, -0.282 and -0.304 respectively, whereas the long-run price elasticities obtained from models (5) and (7) are -0.755 and -0.638 respectively. All the above estimated price elasticities are significant at the 1 percent level, except the ones estimated from equation (5) which are significant at the 8 percent level. These below unity relative price elasticities indicate that food exports of Greece are rather insensitive to relative price changes.

2.2. Export Demand Equations for Beverages and Tobacco (SITC: Section 1)

The share of exported beverages and tobacco in the value of the total exported goods declined from 42.8% in 1954 to 8.6% in 1976, mainly because of the increased participation of the manufactures in the country's exports (chapter II, table 3), and it averaged to 14.6 percent over the whole sample period. Tobacco (unmanufactured), one of the most important export commodities of Greece, makes up more than 85 percent of this group, while the rest consists mainly of alcoholic beverages.

Since the majority of the exported beverages and tobacco is absorbed by the O.E.C.D. countries, the index of total industrial production of the O.E.C.D. countries has been selected as income-proxy. However, Greece exports only unmanufactured cigarette leaf tobacco which is used as input by the tobacco industry of the O.E.C.D. countries. Therefore activity elasticities with respect to the corresponding industrial production have been also estimated. The proxy activity variable used for this purpose is the index of industrial production of food, beverages and tobacco of the O.E.C.D. countries.

The relative price variable introduced into the equation was obtained as the price index of exported beverages and tobacco deflated by a weighted index of domestic prices of the most related commodities (according to the availability of the data) of the seven O.E.C.D. countries which are the main customers of the Greek exports. The particular price series used are the following:

U.S.A. : Price index of food (wholesale prices)

France : Price index of agricultural goods (producer prices)

Germany : Price index of food, beverages and tobacco (producer prices)

Italy : Price index of food (wholesale prices)

Netherlands: Price index of food (wholesale prices)

U.K. : Price index of food (wholesale prices)

Yugoslavia : Price index of agricultural goods (producer prices)

Apart from the above explanatory variables, changes in consumer tastes and technology as well as the anti-smoking campaign may have affected the export demand for beverages and tobacco during the sample period. However, experiments made with a time trend accounting for the effect of these influences did not produce significant results because of the high intercorrelation between the activity and the time variable.

According to the likelihood ratio test, performed in the basic model, the log-linear form was preferred to the linear form. Using seasonally adjusted data, the least squares estimates of the basic equation are as follows:

$$\text{(1)} \quad \log(X_{bt}) = \underset{(1.032)}{1.592} - \underset{(0.173)}{0.103} \log(P^{x}_{bt}/P^{OECD7}_{fag}) + \underset{(0.101)}{+\ 0.704} \log(IP_{OECD})$$

$n=263$, $RSS=65.996$, $R^2=0.178$, $D.W.=1.049$, $X^2(15)=135.839$

where,

X_{bt} = Index of volume of exported beverages and tobacco

P^{x}_{bt} = Price index of exported beverages and tobacco

P^{OECD7}_{fag} = Weighted index of domestic prices of food or agricultural goods of seven importing OECD countries

IP_{OECD} = Index of industrial production of OECD

The price and income elasticities have the expected signs, but the price coefficient is insignificant. Both Durbin-Watson and the $X^2(15)$ Box-Pierce test-statistics are highly significant indicating that first order as well as higher order autocorrelations are present in the residuals and in fact there are several peaks in the residual correlogram. Reestimating (1) subject to an autoregressive error term, a 4th order general autoregressive form was preferred. In this equation, which was accepeted against the unrestricted transformed equation, the price coefficient came out with the wrong sign, but it was insignificant. The same happened with most of the general autoregressive forms which were tried out[1]. After some experimentation we found that the price variable lagged by four months gives better results. So it was decided to revise the basic equation on which the empirical analysis is based, and instead of equation (1) we adopt the following model:

$$(2) \qquad \log(X_{bt}) = \underset{(1.019)}{4.291} - \underset{(0.170)}{0.595} \log(P^{x}_{bt}/P^{OECD7}_{fag})_{t-4} + \underset{(0.099)}{0.618} \log(IP_{OECD})$$

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[1] Almost identical results were obtained, when we used as deflator a weighted index of domestic price series which are the closest approximations to the general price level of the particular OECD countries, i.e.

U.S.A.	: Industrial goods (wholesale prices)
France	: Semi-manufactures (wholesale prices)
Germany	: Total price index (producer prices)
Italy	: Consumer goods (wholesale prices)
Netherlands	: Manufactured goods (wholesale prices)
U.K.	: Total manufacturing output (wholesale prices)
Yugoslavia	: Industrial goods (producer prices)

$n=263$, $RSS=63.120$, $R^2=0.214$[2], $D.W.=1.101$, $X^2=114.987$

Here again, reestimating equation (2) subject to an autoregressive error term, among other forms, a fourth order general autoregressive scheme was preferred. The maximum likelihood estimates are:

$$\text{(3)} \qquad \log(X_{bt}) = \underset{(1.118)}{3.034} - \underset{(0.178)}{0.355} \log(P^{x}_{bt}/P^{OECD7}_{fag})_{t-4} + $$

$$+ \underset{(0.127)}{0.646} \log(IP_{OECD})$$

$$u_t = \underset{(0.062)}{0.434}\, u_{t-1} + \underset{(0.068)}{0.093}\, u_{t-2} - \underset{(0.068)}{0.064}\, u_{t-3} -$$

$$- \underset{(0.062)}{0.138}\, u_{t-4}$$

$n = 263$, $R.S.S. = 48.408$, $X^2(15) = 16.015$

The $X^2(4)$ test statistic on the autocorrelation parameters ρ_1, ρ_2, ρ_3 and ρ_4, gave a value of 69.793 which is highly significant. The income elasticity came up almost with the same value as in (2), whereas the price elasticity decreased in absolute value but remained significant.

The test of the non-linear restrictions imposed by the autoregressive error specification (3) in the corresponding unrestricted transformed equation, which contains one up to four months lagged values of the variables X_{bt} ,

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[2] This low R^2 value is due to the seasonal adjustment of the data series which (mainly the dependent variable) exhibit a strong seasonal variation. If the seasonal dummies are included explicitly into the equation, the R^2 takes the value of 0.722.

$(P^{x}_{bt}/P^{OECD7}_{fag})$ and IP_{OECD}, gave a value of 9.546 for the $X^2(8)$ test statistic, which is not significant; hence, the original equation, autoregressive error specification, is preferred.

Considering now the "common factor" analysis, the general unrestricted dynamic model has the form

$$\text{(4)} \qquad a(L)\log(X_{bt}) = \beta_1(L)\log(P^{x}_{bt}/P^{OECD7}_{fag})_{t-4} + \beta_2(L)\log(IP_{OECD}) + \\ + \text{constant}$$

where the scalar polynomials in the lag operator L are of degree 13. The least squares estimates of (4) gave an equation with 42 coefficients from which only the coefficients $\log(X_{bt})$ lagged one month as well as the coefficient of $\log(P^{x}_{bt}/P^{OECD7}_{fag})_{t-4}$ came out significant and with the correct sign ($n = 259$, R.S.S. = 40.102, $R^2 = 0.490$, D.W. = 2.001, $X^2(15) = 8.570$). This suggests that the model can be simplified and the Wald test for a common factor polynomial confirms this view. The table below gives the Wald criteria as well as their successive differences, for a common factor polynomial of degree m and for m = 11, 12 and 13.

Wald Criteria

m	Wald Criteria	Degrees of Freedom	Differences	Degrees of Freedom
11	1.661	22	-	-
12	4.792	24	3.131	2
13	9.666	26	4.874	2

As can be seen from the above table, both the Wald and the difference Wald criteria take non-significant values indicating that a common factor polynomial of degree 13 exists and therefore a structural equation form without any lags on the variables should be considered and the already estimated original equation, autoregressive error specification (3) is the preferred form.

The above preferred specification (3) shows that the initially extracted set of 13 common roots contains nine zeros which in the general unrestricted dynamic model (4), shorten the lag length by nine for every variable. This was confirmed in an F-test on the coefficients of the variables lagged 5 up to 13 months, which gave a value of 1.158 for the $F(27,206)$ test-statistic which is not significant. Reestimating equation (4) allowing this time the scalar polynomials to be of order 4 and testing for a common factor polynomial we obtained the following values for the Wald criteria.

Wald Criteria

m	Wald Criteria	Degrees of Freedom	Differences	Degrees of Freedom
3	1.719	6	-	-
4	3.595	8	1.876	2

Here again, it seems fairly clear that on the basis of both the Wald and the difference Wald criteria, the hypothesis that m = 4 is accepted and the original equation , autoregressive error specification (3) should be considered as the data generation process.

Turning now to a test of the seasonal adjustment of the data , we reestimated equation (3) using unadjusted data. The F-test constructed from the corresponding residual sums of squares gave a value of 17.161 for the $F(11,245)$ test-statistic which is highly significant. Finally, according to the likelihood ratio test, performed in the preferred equation (3), the log-linear form is preferred to the linear form.

The estimated activity elasticity of 0.646 in equation (3) shows that exports of beverages and tobacco are income inelastic with respect to the index of industrial production of the O.E.C.D. countries. Reestimating equation (3) employing this time as activity variable the index of industrial production of food, beverages and tobacco of the O.E.C.D. countries, we obtained an activity elasticity of 0.829[3]. This activity elasticity although higher in value than the one before, is still below unity indicating that the foreign demand for the Greek exports of beverages and tobacco is income inelastic. However, a more appropriate activity variable (e.g. index of cigarette production) could produce a higher activity elasticity, since, as it mentioned in the beginning of the section, Greece exports only cigarette leaf tobacco. Moreover, the

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[3] Here again, a 4th order general autoregressive form was preferred and the autoregressive error hypothesis was accepted against the unrestricted transformed equation ($X^2(8) = 14.369$).

tobacco which is exported from Greece is of the oriental type mainly used in cigarette blends, the American type of which has dominated the world market.

The estimated relative price elasticity of -0.355, though significant, is low enough indicating that changes in relative prices exert a small influence on the export demand for beverages and tobacco.

2.3. Export Demand Equations for Raw Materials (SITC: Section 2)

For the period under review exports of raw materials accounted for about 13 percent in the value of the total exports of goods. This group of exported commodities mainly includes raw cotton, crude minerals, ores of base metals, and hides and skins, whose shares averaged to 28.1, 22.7, 21.4 and 16.3 percent respectively for the period 1954 - 1976.

The bulk of the exported raw materials of Greece is destined for the O.E.C.D. countries, and is used by the importing countries as input to their industrial sectors. Therefore, the level of industrial production of the O.E.C.D. countries appears to be the most suitable activity variable.

To obtain the relative price variable, the price index of exported raw materials was deflated by a weighted index of domestic prices of raw materials (or the most relevant commodities according to data availability) of seven O.E.C.D. countries which are the major buyers of Greek exports. The price series used are the following:

U.S.A. : Price index of industrial goods (wholesale prices)

France : Price index of raw materials (wholesale prices)

Germany : Price index of intermediate goods (producer prices)

Italy : Price index of industrial materials (wholesale prices)

Netherlands: Price index of intermediate goods (wholesale prices)

U.K. : Price index of raw materials (wholesale prices)

Yugoslavia : Price index of industrial goods (producer prices)

Sargan's maximum likelihood ratio gave a value of 1.19, for the basic equation, which indicates that the log-linear form is preferred to the linear form. The least squares estimates of the basic model, from seasonally adjusted data, are as follows:

$$(1) \qquad \log(X_{rm}) = \underset{(0.703)}{5.128} - \underset{(0.163)}{1.052} \log(P^{x}_{rm}/P^{OECD7}_{rm}) + \underset{(0.051)}{0.934} \log(IP_{OECD})$$

$n=263$, $RSS=15.341$, $R^2=0.568$, $D.W.=1.270$, $X^2(15)=75.583$

where,

X_{rm} = Index of volume of exported raw materials

P^{x}_{rm} = Price index of exported raw materials

P_{rm}^{OECD7} = Weighted index of domestic prices of raw materials of seven importing OECD countries

IP_{OECD} = Index of industrial production of OECD .

The estimated regression coefficients are highly significant and they have the theoretically expected signs. Both Durbin-Watson and the $X^2(15)$ Box-Pierce test statistics are significant, and in fact there are autocorrelations biger than twice their standard errors in the residual correlogram at lags 1, 2, 3 and 4. Reestimating equation (1) subject to a simple ℓth order autoregressive error, we obtained, as it was expected, a significant autocorrelation parameter for ℓ = 1, 2, 3 and 4. In all four cases the original equation, autoregressive error specification was accepted against the unrestricted transformed equation, and the estimated activity and price elasticities were very close to the ones of equation (1). However, a general ℓth order autoregressive scheme was preferred and in likelihood ratio tests the hypothesis that $\ell = 2$ was accepted against other values of ℓ.

The maximum likelihood estimates are:

$$(2) \qquad \log(X_{rm}) = \underset{(0.902)}{5.528} - \underset{(0.202)}{1.126} \log(P_{rm}^{x}/P_{rm}^{OECD7}) + \underset{(0.084)}{0.923} \log(IP_{OECD})$$

$$u_t = \underset{(0.061)}{0.310}\, u_{t-1} + \underset{(0.062)}{0.153}\, u_{t-2} + \varepsilon_t$$

$n = 263$, R.S.S. = 12.967, $X^2(15) = 16.619$

The test of the non-linear restrictions, imposed in the estimation of the general autoregressive error specification (2), gave a value of 4.641 for the $X^2(4)$ test-statistic, which is not significant; hence, the original equation, autoregressive error specification is preferred.

Considering now the "common factor" analysis we begin with the general unrestricted dynamic model

$$(3) \qquad a(L)\log(X_{rm}) = \beta_1(L)\log(P^x_{rm}/P^{OECD7}_{rm}) + \beta_2(L)\log(IP_{OECD}) + \text{constant}$$

where the scalar polynomials in the lag operator L are of degree 13. The least squares estimates of (3) gave an equation with 42 coefficients from which only the coefficients of $\log(X_{rm})$, lagged one and two months, $\log(IP_{OECD})$, lagged one and twelve months, and $\log(P^x_{rm}/P^{OECD7}_{rm})_t$ were significantly different from zero (n = 263, R.S.S. = 11.138, R^2 = 0.686, D.W. = 1.985, $X^2(15)$ = 6.161). The fact that only five of the forty two estimated coefficients are significantly different from zero suggests that some simplification of the model is possible and the tests for common factor polynomials confirm this view. The table below gives the Wald criteria, as well as their successive differences, for a common factor polynomial of degree m and for m = 10, 11 , 12 and 13.

Wald Criteria

m	Wald Criterion	Degrees of Freedom	Differences	Degrees of Freedom
10	1.632	20	-	-
11	2.784	22	1.152	2
12	9.015	24	6.230*	2
13	17.997	26	8.982*	2

* Significant at the 5% level

Taking first the Wald criteria, none of them takes a significant value, and therefore the hypothesis that m = 13 should be accepted. This suggests that we require a structural equation form without any lags on the variables and the already estimated specification (2) is the preferred form.

Taking now the Wald difference criteria, it appears from the above table that on the basis of the difference between m = 12 and m = 13, and between m = 11 and m = 12, the hypothesis that m = 13 or 12 might be regected in favor of m = 11. This indicates that we require a structural equation form in which at least some of the variables of equation (1) have two extra lags. Reestimating the basic equation including this time all variables lagged one and two months, only the lagged values of the dependent variable produced a significant coefficient, and on estimation subject to an autoregressive error term we did not find any significant autocorrelation (as it was expected in view of (2)). So, after dropping some non-significant variables, we obtained the following equation:

$$\text{(4)} \quad \log(X_{rm}) = \underset{(0.687)}{3.809} + \underset{(0.061)}{0.305} \log(X_{rm})_{t-1} + \underset{(0.058)}{0.096} \log(X_{rm})_{t-2} - \underset{(0.157)}{0.799} \log(P^{x}_{rm}/P^{OECD7}_{rm}) + \underset{(0.075)}{0.572} \log(IP_{OECD})$$

$n=263$, $RSS=13.221$, $R^2=0.628$, $D.W.=1.960$, $\chi^2(15)=14.754$

which according to the difference Wald criteria is the preferred form. Notice, however, that the long-run income and price elasticities obtained from the above model are very close to those estimated by the static model (2).

The above preferred autoregressive error specification (2) shows that the initially extracted set of 13 common roots (according to the Wald criteria, or 11 common roots according to the difference Wald criteria) contains eleven zeros which, in the general unrestricted dynamic model (4), shorten the lag length by eleven for every variable. This was confirmed in an F-test on the coefficients of the variables lagged 3 up to 13 months, which gave a value of 0.916 for the F(33,210) test-statistic which is not significant. Reestimating (4) allowing this time the scalar polynomials to be of order 2 and testing for a common factor polynomial we obtained the following values of the Wald criteria.

Wald Criteria

m	Wald Criteria	Degrees of Freedom	Diffe-rences	Degrees of Freedom
1	1.343	2	1.343	2
2	2.385	4	1.042	2

As can be seen from the above table, none of the Wald or the difference Wald criteria takes a significant value. So, the common factor analysis as applied in the simplified version of (4), leads us to accept a structural equation form without any lags on the variables and with an autoregressive error term, i.e. model (2), as the preferred form.

To test the seasonal adjustment of the data, we reestimated equation (2) using unadjusted data. The F-test constructed from the corresponding residual sums of squares gave a value of 14.670 for the F(11,247) test-statistic which is highly significant. Finally, according to the likelihood ratio test performed in model (2), the log-linear form is preferred to the linear form.

The estimated activity elasticity of 0.923[1] in equation (2) is highly significant, while its size slightly below unity is in agreement with a priori expectations. The volume of exported raw materials is expected to change in much the same proportion as the level of industrial production of the importing countries and the above activity elasticity confirms this proportionality hypothesis.

On the other hand, the price elasticity estimated from equation (2) amounts to -1.126, and is highly significant. This above unity price elasticity indicates that Greek exports

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[1] Experiments made with the index of industrial production of the manufacturing industries only, of the OECD countries, produced approximately the same results.

of raw materials face noticeable price competition by the raw materials produced by the O.E.C.D. member countries.

2.4. Export Demand Equations for Chemicals (SITC: Section 5)

Exports of chemicals averaged 4.9 percent of the value of total exports of goods during the period 1954 - 1976. In the early years of the sample, exports of chemicals were low and stagnant, whereas in later years they started picking up. This is due to the construction and operation, in the early sixties, of several large plants producing among other things, polysterene, nitrogenous fertilizers, phosphates, aluminium oxide, basic chemicals, fertilizers, petrochemicals and ammonia. In view of the above, exports of chemicals increased at an annual rate of 1.4 percent in the years 1954 - 1966, 128.6 percent in 1966 - 1967, 105.3 percent in 1967 - 1968, 8.4 percent in 1968 - 1976, and 11.3 percent during the whole sample period.

Raw chemical materials, which are utilised by the chemical industries of the importing countries, make up more than 85 percent of this group, while the rest consists of chemical manufactures such as pharmaceutical products mainly, and cosmetics. Therefore, the index of chemical production of the OECD countries has been selected to represent the activity variable[1] entering the export demand equation for chemicals.

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[1] The proxy activity variable used is the index of industrial production of chemicals, rubber, and petroleum and coal products of the OECD countries.

The relative price variable introduced into the equation was obtained as the price index of exported chemicals deflated by a weighted index of domestic prices of chemicals (or the most related commodoties when the relevant price index was not available) of seven OECD countries which absorb more than 70 percent of Greece's exports of chemicals. The particular price series used are the following:

U.S.A.	: Price index of chemicals (wholesale prices)
France	: Price index of chemicals (wholesale prices)
Germany	: Price index of intermediate goods (producer prices)
Italy	: Price index of chemicals, fuel and lubricants (wholesale prices)
Netherlands	: Price index of chemicals (producer prices)
U.K.	: Price index of chemicals (wholesale prices)
Yugoslavia	: Price index of industrial goods (producer prices)

Sargan's likelihood ratio test, performed in the basic model, gave a value of 2.096 which indicates that the log-linear form is preferred to the linear form. Using seasonally adjusted data, the least-squares estimates of the basic equation are as follows:

$$(1) \qquad \log(X_{ch}) = \underset{(0.919)}{-\,2.224} - \underset{(0.244)}{0.192}\log(P^{x}_{ch}/P^{OECD7}_{ch}) + \underset{(0.092)}{1.601}\log(ICP_{OECD})$$

$n=263$, $RSS=91.963$, $R^2=0.637$, $D.W.=0.903$, $X^2(15)=876.408$

where,

X_{ch} = Index of volume of exported chemicals

P^x_{ch} = Price index of exported chemicals

P^{OECD7}_{ch} = Weighted index of domestic prices of chemicals of seven importing OECD countries

ICP_{OECD} = Index of industrial production of chemicals of OECD

The estimated price and income elasticities have the expected signs, but the price elasticity is insignificant. Both the Durbin-Watson and the $X^2(15)$ Box-Pierce test-statistics are highly significant indicating that first order as well as higher order autocorrelations are present in the residuals. In fact there are autocorrelations in the residual correlogram at all lags, taking values in the range 0.39 to 0.55.

The above indicate that a general autoregressive error scheme should be considered. So, we reestimated equation (1) allowing the error term to follow a general ℓth order autoregression which finally was preferred to a simple autoregressive form, and in likelihood ratio tests the hypothesis that $\ell = 6$ was accepted against other alternative values.

The maximum likelihood estimates are:

(2) $$\log(X_{ch}) = \underset{(1.901)}{0.775} - \underset{(0.237)}{1.182} \log(P^x_{ch}/P^{OECD7}_{ch}) + \underset{(0.402)}{1.943} \log(ICP_{OECD})$$

$$u_t = \underset{(0.063)}{0.236}\, u_{t-1} + \underset{(0.064)}{0.162}\, u_{t-2} + \underset{(0.065)}{0.094}\, u_{t-3} + \underset{(0.064)}{0.139}\, u_{t-4} + \underset{(0.064)}{0.113}\, u_{t-5} + \underset{(0.063)}{0.125}\, u_{t-6}$$

$n = 263$, R.S.S. $= 49.148$, $X^2(15) = 18.168$

The $X^2(6)$ test-statistic on the autocorrelation parameters ρ_i , $i = 1, \ldots, 6$, gave a value of 176.173 which is highly significant. The standard errors of the estimated coefficients have increased, indicating that they were initially underestimated in the presence of autocorrelated errors. Also, both the price and income coefficients increased in absolute value, and the price coefficient became highly significant.

The test of the non-linear restrictions imposed by the autoregressive error specification (2) in the corresponding unrestricted transformed equation, which contains one up to six months lagged values cf all three variables, gave a value of 11.393 for the $X^2(12)$ test statistic, which is not significant; hence, the original equation, autoregressive error specification, is preferred.

Following now the "common factor" analysis, the general unrestricted dynamic model has the form

$$\text{(3)} \qquad \log(X_{ch}) = \beta_1(L)\log(P^x_{ch}/P^{OECD7}_{ch}) + \beta_2(L)\log(ICP_{OECD}) + \\ + \text{constant}$$

where the scalar polynomials in the lag operator L are of order 13. The least squares estimates of (5) gave an equation with 42 coefficients, most cf which are insignificant (and with wrong sign) because of the collinearity between the different lags of the variables ($n = 263$, R.S.S. $= 42.841$, $R^2 = 0.831$, D.W. $= 1.954$, $X^2(15) = 7.098$). In particular, only the coefficient of the dependent variable lagged one month as well as the coefficient of $\log(P^x_{ch}/P^{OECD7}_{ch})_t$ came up significant (and with the expected sign). So, with only two of the forty two coefficients significant, a simplification of the model is possible and this is confirmed by the tests for common factor polynomials.

The table below gives the Wald criteria, as well as their successive differences, for a common factor polynomial of degree m and for m = 11, 12 and 13.

Wald Criteria

m	Wald Criteria	Degrees of Freedom	Diffe- rences	Degrees of Freedom
11	0.795	22	-	-
12	3.889	24	3.094	2
13	6.515	26	2.626	2

As can be seen from the above table, neither the Wald nor the difference Wald criteria take significant values and therefore the hypothesis that m = 13 is accepted. This suggests that we should consider a structural equation form without any lags on the variables and an autoregressive error specification, at least in the first instance , of general 13th order form. When this model was estimated, the hypothesis that $\rho_7 = \ldots = \rho_{13} = 0$ was accepted ($X^2(7) = 11.086$), as already mentioned, and therefore model (2) is the preferred form.

The non-significant values of $\rho_7, \ldots, \rho_{13}$ in the general autoregressive error specification, indicate that the common factor polynomial of order 13 contains seven roots equal to zero which, in the general unrestricted dynamic model (3), shorten the lag length by seven for every variable. To verify that we reestimated the general unrestricted equation (3) allowing this time the scalar polynomials in the lag operator L to be of order 6 (n = 263, R.S.S. = 47.065, R^2 = 0.814, D.W. = 2.002, $X^2(15)$ = 19.057). The F-test between the two general unrestricted forms gave a value of 0.986 for the F(21,210) test-statistic which is not significant. Testing now for a common factor polynomial in the simplified version of (3) we obtained the following values for the Wald and difference Wald criteria.

Wald Criteria

m	Wald Criteria	Degrees of Freedom	Differences	Degrees of Freedom
5	1.789	10	-	-
6	3.199	12	1.410	2

Here again, none of the Wald or the difference Wald criteria takes a significant value and the hypothesis that m = 6 is accepted suggesting that model (2) is our data generation process.

The above empirical analysis indicates that the overall dynamics in the export demand equation for chemicals is of 6th order. Both the model selection procedure which starts with the simplest model and the alternative procedure which begins with a general unrestricted model, lead to a specification where the overall dynamics in the equation is expressed only through error dynamics.

Turning now to a test of the seasonal adjustment of the data, we reestimated equation (2) using unadjusted data. The F-test constructed from the corresponding residual sums of squares gave a value of 4.643 for the $F(11,243)$ test statistic which is highly significant. Finally, according to the likelihood ratio test, performed in the preferred equation (2), the log-linear form is preferred to the linear form.

The estimated activity elasticity of 1.943 shows that exports of chemicals are income elastic with respect to the index of chemical production of the O.E.C.D. countries. This rather high activity elasticity is partly attributed to the structural changes mentioned in the beginning of the section, namely the rapid growth of the Greek chemical industry resulting in an increasing supply of the commodities in question.

From the behaviour of the residuals it becomes evident that the above structural changes took place at the end of 1967. So, a dummy variable taking the value one for the period 1968 - 1976 and zero outside it, was introduced into our equation and the above empirical analysis was repeated. Here again a 6th order general autoregressive error form was preferred and the autoregressive error hypothesis was accepted against the unrestricted transformed equation ($X^2(12) = 18.019$). Also, the "common factor" analysis gave similar results to the ones previously described. The new estimated equation is:

$$(4) \qquad \log(X_{ch}) = \underset{(1.826)}{3.122} - \underset{(0.237)}{1.219} \log(P^{x}_{ch}/P^{OECD7}_{ch}) +$$

$$+ \underset{(0.392)}{1.360} \log(ICP_{OECD}) + \underset{(0.303)}{0.684} D$$

$$u_t = \underset{(0.063)}{0.209} u_{t-1} + \underset{(0.064)}{0.150} u_{t-2} + \underset{(0.065)}{0.076} u_{t-3} +$$

$$+ \underset{(0.065)}{0.123} u_{t-4} + \underset{(0.065)}{0.112} u_{t-5} + \underset{(0.063)}{0.129} u_{t-6}$$

$$n = 263, \text{ R.S.S.} = 48.368, \ X^2(15) = 18.815$$

where,

D - 1 for 1968 - 1976 and 0 elsewhere

It appears then that the activity elasticity is reduced to 1.36[2] which is reasonable.

On the other hand the estimated price elasticity of -1.219, being highly significant, reflects the price competition that Greek exportable chemicals face in the foreign markets

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[2] When the total index of industrial production of the OECD countries was employed as an activity variable, the estimated income elasticity was 1.818.

2.5. Export Demand Equations for Manufactured Goods (SITC: Section 6)

The share of exported manufactures in the value of the total exports of goods increased from 5.4% in 1954 to 31.7% in 1976, and it averaged 25.6 percent over the whole sample period. This noticeable increase of the share of manufactures in total exports of goods is attributed to the rapid growth of Greek manufacturing, mainly during the decade 1960-1970, and particularly of the textile, nonmetallic minerals, metal and machinery industries. In view of the above, exports of manufactures increased at an average annual rate of 6.1 percent in the years 1954 - 1961, 28.7 percent in 1961 - 1976 (for the period 1961 - 70 the annual rate of growth was 37.2 percent), and 21.0 percent during the whole sample period. This export group mainly includes textile products (mainly yarns and threads), lime, cement etc., and metals which account for more than 60 percent of the exports of manufactures. Also, aluminium, exported for the first time in 1966, has become one of the leading commodities in the exports of manufactures, accounting for about 20 percent of the group in question.

Since the bulk of the exported manufactures of Greece is destined for the O.E.C.D. countries, the total index of industrial production of the O.E.C.D. countries has been selected as an income-proxy.

The relative price variable introduced into the equation was obtained as the price index of exported manufactures deflated by a weighted index of domestic prices of manufactures (or the most related commodities according to the availability of the data) of the seven O.E.C.D. countries which are the main customers of the Greek exports[1]. The particular price series used are the following:

U.S.A.	: Price index of industrial goods (wholesale prices)
France	: Price index of semi-manufactures (wholesale prices)
Germany	: Price index of consumer goods (producer prices)
Italy	: Price index of consumer (non-food) goods (wholesale prices)
Netherlands	: Price index of manufactures (wholesale prices)
U.K.	: Price index of manufactures (wholesale prices)
Yugoslavia	: Price index of industrial goods (producer prices)

For the basic model, Sargan's maximum likelihood ratio gave a value of 5.39 which clearly indicates that the log-linear form is preferred to the linear form. The least-squares estimates of the basic equation obtained from seasonally adjusted data, are as follows:

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[1] A dummy variable, taking the value 1 in 1966-76 and 0 elsewhere, was introduced into the equation to take account of the addition of aluminium in the exports of manufactures in the period 1966-76, but it did not produce significant results.

$$(1) \qquad \log(X_{mf}) = -\underset{(1.064)}{21.558} + \underset{(0.166)}{0.572} \log(P^{x}_{mf}/P^{OECD7}_{mf}) + $$

$$+ \underset{(0.086)}{5.068} \log(IP_{OECD})$$

n=263, RSS=29.963, R^2=0.956, D.W.=1.160, $X^2(15)$=237.663

where,

X_{mf} = Index of volume of exported manufactures

P^{x}_{mf} = Price index of exported manufactures

P^{OECD7}_{mf} = Weighted index of domestic prices of manufactures of seven importing OECD countries

IP_{OECD} = Index of industrial production of OECD

It appears from the estimated equation that the price and income coefficients are significant, but the price coefficient has the wrong sign. This is due to the presence of intercorrelation between the income and price variables (r = -0.655, whereas the correlation coefficient between exports and prices is -0.606 and between exports and income 0.977). Experiments made with lagged values of the price and income variables produced similar results because in all cases the correlation coefficient between price and income was biger in absolute value than the correlation coefficient between exports and price.

Both Durbin-Watson and the $X^2(15)$ Box-Pierce test-statistics are highly significant indicating that first order as well as higher order autocorrelations are present in the residuals. Reestimating equation (1) subject to an autoregressive error term, a general ℓth order autoregression was preferred to a simple autoregressive form, and in likelihood ratio tests the hypothesis that $\ell = 7$ was accepted against other alternative values.

The maximum likelihood estimates are:

$$(2) \qquad \log(X_{mf}) = \underset{(2.383)}{-\,11.016} - \underset{(0.211)}{0.586}\log(P^{x}_{mf}/P^{OECD7}_{mf}) + \underset{(0.475)}{3.918}\log(IP_{OECD})$$

$$u_t = \underset{(0.062)}{0.220}\,u_{t-1} + \underset{(0.064)}{0.187}\,u_{t-2} + \underset{(0.064)}{0.091}\,u_{t-3} + \underset{(0.063)}{0.161}\,u_{t-4} + \underset{(0.062)}{0.009}\,u_{t-5} + \underset{(0.060)}{0.090}\,u_{t-6} + \underset{(0.057)}{0.148}\,u_{t-7}$$

$n = 263$, R.S.S. $= 18.664$, $X^2(15) = 9.934$

We notice that the estimated income elasticity decreased in value and the price coefficient came out significant and with the correct sign.

Turning now to a test of the non-linear restrictions imposed in estimation, we find that the original specification, autoregressive error hypothesis (2), is rejected in favor of the unrestricted transformed equation ($X^2(14) = 44.767$). This equation, containing one up to seven months lagged values of all three variables and an independent error term, achieves a significantly lower residual sum of squares than the restricted form given above, and after dropping some non-significant variables we obtain

$$(3) \qquad \log(X_{mf}) = \underset{(1.529)}{-\,2.348} + \underset{(0.060)}{0.200}\log(X_{mf})_{t-1} + \underset{(0.061)}{0.164}\log(X_{mf})_{t-2} + \underset{(0.060)}{0.154}\log(X_{mf})_{t-4} + \underset{(0.056)}{0.111}\log(X_{mf})_{t-6} + \underset{(0.054)}{0.185}\log(X_{mf})_{t-7} -$$

$$- \underset{(0.131)}{0.161} \log(P^{x}_{mf}/P^{OECD7}_{mf}) + \underset{(0.316)}{0.867} \log(IP_{OECD})_{t-7}$$

$n=263$, $RSS=17.795$, $R^2=0.974$, $D.W.=2.039$, $X^2(15)=13.918$

As mentioned before, due to the intercorrelation between the income and price variables, we were not able to obtain a significant price coefficient. The above equation (with a non-significant price coefficient) was the best one after all combinations between $\log(P^{x}_{mf}/P^{OECD7}_{mf})_{t-i}$ and $\log(IP_{OECD})_{t-j}$, for $i,j = 0, 1, \ldots, 7$, were tried out. However, after some more experimentation, we found that the price and income variables lagged by thirteen months resulted in a significant price coefficient with the correct sign. The new estimated equation is:

$$\text{(4)} \quad \log(X_{mf}) = - \underset{(1.638)}{1.371} + \underset{(0.061)}{0.176} \log(X_{mf})_{t-1} +$$

$$+ \underset{(0.061)}{0.151} \log(X_{mf})_{t-2} + \underset{(0.060)}{0.148} \log(X_{mf})_{t-4} +$$

$$+ \underset{(0.055)}{0.126} \log(X_{mf})_{t-6} + \underset{(0.053)}{0.195} \log(X_{mf})_{t-7} -$$

$$- \underset{(0.148)}{0.343} \log(P^{x}_{mf}/P^{OECD7}_{mf})_{t-13} + \underset{(0.332)}{0.860} \log(IP_{OECD})_{t-13}$$

$n=263$, $RSS=17.417$, $R^2=0.974$, $D.W.=2.044$, $X^2(15)=12.086$

All the coefficients came up almost with the same values as in (3) with an exception the price coefficient which increased in absolute value and became significant. So, equation (4) is the preferred form.

Considering now the "common factor" analysis the general unrestricted dynamic model has the form

$$\text{(5)} \quad a(L)\log(X_{mf}) = \beta_1(L)\log(P^{x}_{mf}/P^{OECD7}_{mf}) + \beta_2(L)\log(IP_{OECD}) +$$

$$+ \text{constant}$$

where the scalar polynomials in the lag operator L are of order 13. The least squares estimates of (5) gave an equation with 42 coefficients from which only the coefficient of $\log(X_{mf})$ lagged seven months as well as the coefficients of $\log(P^{X}_{mf}/P^{OECD7}_{mf})$ at time t and t-13 came out significant and with the correct sign (n = 263, R.S.S. = 14.130, $R^2 = 0.979$, D.W. = 1.984, $X^2(15) = 9.517$).

The Wald test for a common factor polynomial of degree 13, gives a value of 22.984 for the $X^2(26)$ test-statistic which is not significant. This suggests that a structural equation form without any lags on the variables and a general autoregressive error specification, i.e. model (2), should ce considered as the data generation process.

The above specification (2), preferred among other autoregressive forms, shows that the initially extracted set of 13 common roots contains six zeros which in the general unrestricted dynamic model (5), shorten the lag length by six for every variable. This was confirmed in an F-test on the coefficients of the variables lagged 8 up to 13 months, which gave a value of 1.332 for the F(18,210) test-statistic which is not significant. Reestimating equation (5) allowing this time the scalar polynomials to be of order 7 and testing for a common factor polynomial of degree m, we obtained the following values for the Wald criteria.

Wald Criteria

m	Wald Criteria	Degrees of Freedom	Diffe-rences	Degrees of Freedom
5	1.115	10	-	-
6	3.518	12	2.403	2
7	15.009	14	11.491*	2

*Significant at the 1% level.

Taking first the Wald criteria, none of them takes a significant value and therefore the hypothesis that m = 7 is accepted. This suggests that a structural equation form without any lags on the variables and an autoregressive error specification of general 7th order form should be considered, and the already estimated model (2) is the preferred form. Taking now the Wald difference criteria, it appears from the above table that, on the basis of the difference between m = 7 and m = 6, the hypothesis that m = 7 might be rejected in favor of m = 6. This indicates that we require a structural equation form in which at least some of the variables of equation (1) have one extra lag. However, experiments made with one period lagged values of the variables and general autoregressive error forms, failed to produce acceptable results because of different local minima (simple autoregressive forms were inadequate to describe the data generation process).

The above empirical analysis indicates that the overall dynamics in the export demand equation for manufactures is of

7th order. The model selection procedure which starts from the static form of the basic equation leads to a model where the overall dynamics in the equation is expressed only through systematic dynamics (models (3) and (4)). On the contrary, starting from the general unrestricted dynamic model, the common factor analysis based on the Wald criteria suggests that the overall dynamics in the equation is expressed only through error dynamics (model (2)).

Turning now to a test of the seasonal adjustment of the data, we reestimated both equations (2) and (4) using unadjusted data. The F-tests constructed from the corresponding residual sums of squares gave the values of 2.60 and 2.807 for the $F(11,242)$ and $F(11,244)$ test-statistics respectively, which are significant. Finally, according to the likelihood ratio test performed in both the above models, the log-linear form is preferred to the linear form.

The short-run income elasticities estimated from models (2) and (4) are 3.918 and 0.860 respectively, whereas the long-run income elasticity obtained from model (4) is 4.208. It appears then that with respect to the index of industrial production of the O.E.C.D. countries, the Greek exports of manufactures are income elastic. However, these income elasticities are higher than what we would expect and they are partly attributed to the structural changes mentioned in the beginning of the section, i.e. the rapid growth of the Greek manufacturing during the sample period.

With respect to the price variable, the estimated short-run elasticities from models (2) and (4) are -0.586 and -0.343

respectively, whereas the long-run price elasticity obtained from model (4) is -1.677, all being significant. These price elasticities, and particularly the last one, indicate the price competition that Greek exports of manufactures face in the international market.

2.6. Export Demand Equations for Total Merchandise Exports (SITC: Sections 0 - 9)

Thus far, we have estimated the demand equations for the major groups of exported commodities which contribute about 85 percent in the value of total exported goods. In this section we attempt to estimate a demand equation for total Greek exports.

It was mentioned in the preceding sections that the O.E.C.D. countries are the main source of demand for Greek exports. Therefore, the index of total industrial production of the O.E.C.D. countries has been selected as income-proxy. To obtain the relative price variable, the price index of the exported commodities was deflated by a weighted index of domestic prices which are the closest approximations to the general price level of seven O.E.C.D. countries which are the main customers of the Greek exports. The particular price series used are the following:

U.S.A. : Price index of industrial goods (wholesale prices)

France : Price index of semi-manufactures (wholesale prices)

Germany : Total price index (producer prices)

Italy : Price index of consumer goods (wholesale prices)

Netherlands: Price index of manufactures (wholesale prices)

U.K. : Total price index of manufacturing output (wholesale prices)

Yugoslavia : Price index of industrial goods (producer prices)

According to the likelihood ratio test, performed in the basic model, the log-linear form was preferred to the linear form. Using seasonally adjusted data, the least squares estimates of the basic equation are as follows:

$$(1) \qquad \log(X_T) = \underset{(0.828)}{-\;10.430} + \underset{(0.162)}{1.154} \log(P_T^x / P_d^{OECD7}) + \underset{(0.049)}{2.103} \log(IP_{OECD})$$

$n=263$, $RSS=16.217$, $R^2=0.878$, $D.W.=0.820$, $X^2(15)=380.596$

where,

X_T = Index of volume of total exports

P_T^x = Price index of total exports

P_d^{OECD7} = Weighted index of domestic prices of seven O.E.C.D. countries

IP_{OECD} = Index of industrial production of O.E.C.D.

Although exports are significantly correlated with relative prices taken alone ($r = -0.2$), due to the presence of intercorrelation between the income and price variables ($r = -0.3$), the price coefficient came out with the wrong sign when both variables were included in the equation. Experiments made with lagged values of the price and income variables produced similar results for the same reason.

Both Durbin-Watson and the $X^2(15)$ Box-Pierce test statistics are highly significant indicating that first order as well as higher order autocorrelations are present in the residuals. Reestimating equation (1) subject to an autoregressive error term, a general ℓth order autoregression was preferred to a simple autoregressive form, and in likelihood ratio tests the hypothesis that $\ell = 12$ was accepted against other alternative values.

The maximum likelihood estimates are:

$$(2) \qquad \log(X_T) = \underset{(1.715)}{-\,4.184} + \underset{(0.194)}{0.191} \log(P_T^x / P_d^{OECD7}) + \underset{(0.311)}{1.733} \log(IP_{OECD})$$

$$u_t = \underset{(0.062)}{0.353}\, u_{t-1} + \underset{(0.064)}{0.082}\, u_{t-2} + \underset{(0.064)}{0.007}\, u_{t-3} + \underset{(0.064)}{0.019}\, u_{t-4} - \underset{(0.064)}{0.028}\, u_{t-5} - \underset{(0.063)}{0.003}\, u_{t-6} - \underset{(0.063)}{0.045}\, u_{t-7} + \underset{(0.064)}{0.025}\, u_{t-8} - \underset{(0.063)}{0.028}\, u_{t-9} + \underset{(0.064)}{0.071}\, u_{t-10} + \underset{(0.064)}{0.222}\, u_{t-11} + \underset{(0.062)}{0.250}\, u_{t-12}$$

$n = 263$, R.S.S. $= 7.747$, $X^2(15) = 10.715$

We notice that the price coefficient came out again with the wrong sign but it is insignificant.

Due to the highly significant values of the first and twelveth autoregressive parameters ρ_1 and ρ_{12}, the factorised error specification $(1-\rho_1 L)(1-\rho_{12} L^{12})u_t = \varepsilon_t$ was considered and though accepted against the unrestricted error specification $u_t = \rho_1 u_{t-1} + \rho_{12} u_{t-12} + \rho_{13} u_{t-13} + \varepsilon_t$ ($X^2(1) = 0.44$), the latter was rejected in the first place in favor of the general 13th order autoregressive form ($X^2(10) = 88.624$) (more generally, in the general 13th order autoregression $u_t = \sum_{i=1}^{13} \rho_i u_{t-i} + \varepsilon_t$, the joint hypothesis $\rho_2 = \dots = \rho_{11} = 0$, $\rho_{13} = -\rho_1 \rho_{12}$ was rejected - $X^2(11) = 88.744$).

Turning now to a test of the non-linear restrictions imposed in estimation, we find that the original specification, autoregressive error hypothesis (2) is rejected in favor of the unrestricted transformed equation ($X^2(24) = 62.028$). This equation, containing one up to twelve months lagged values of all three variables and an independent error term, achieves a substantially lower residual sum of squares than the restricted form given above, and after dropping some non-significant variables we obtain

$$(3) \qquad \log(X_T) = \underset{(0.659)}{-\,0.038} + \underset{(0.050)}{0.330} \log(X_T)_{t-1} + \underset{(0.050)}{0.258} \log(X_T)_{t-11} + \underset{(0.058)}{0.335} \log(X_T)_{t-12} - \underset{(0.113)}{0.102} \log(P_T^x/P_d^{OECD7})_{t-4} + \underset{(0.101)}{0.208} \log(IP_{OECD})_{t-11}$$

$n=263$, $RSS=6.956$, $R^2=0.948$, $D.W.=2.166$, $X^2(15)=14.839$

The above equation was the best one after all combinations between $\log(P_T^x/P_d^{OECD7})_{t-i}$ and $\log(IP_{OECD})_{t-j}$, for $i,j = 0, 1, \dots, 12$, were tried out. As mentioned before, due to the intercorrelation between the income and price variables, in most of these combinations the price coefficient came out with the wrong sign but it was insignificant.

Following now the "common factor" analysis we begin with the general unrestricted dynamic model

$$(4) \qquad a(L)\log(X_T) = \beta_1(L)\log(P_T^x/P_d^{OECD7}) + \beta_2(L)\log(IP_{OECD}) + \\ + \text{constant}$$

where the scalar polynomials in the lag operator L are of order 13. The least squares estimates of (4) gave an equation with 42 coefficients from which only the coefficients of the dependent variable lagged one, eleven and twelve months came out significant (n = 263, R.S.S. = 6.088, R^2 = 0.954, D.W. = 1.998, $X^2(15)$ = 10.085). Testing for a common factor polynomial of degree m, and for m = 11, 12 and 13, we obtained the following Wald and difference Wald criteria.

Wald Criteria

m	Wald Criterion	Degrees of Freedom	Diffe-rences	Degrees of Freedom
11	1.040	22	-	-
12	1.986	24	0.946	2
13	11.171	26	9.185*	2

*Significant at the 2% level

Taking first the Wald criteria, none of them takes a significant value and therefore the hypothesis that m = 13 is accepted. This suggests that a structural equation form without any lags on the variables and a general autoregressive error specification, i.e. model (2), should be considered as the preferred form.

Taking now the Wald difference criteria, it appears from the above table that, on the basis of the difference between $m = 12$ and $m = 13$, the hypothesis that $m = 13$ might be rejected in favor of $m = 12$. This indicates that we require a structural equation form in which at least some of the variables of equation (1) have one extra lag. However, experiments made with one period lagged values of the variables and general autoregressive error forms , failed to produce acceptable results because of different local minima (simple autoregressive forms were inadequate to describe the data generation process).

The above specification (2), preferred among other autoregressive forms, shows that the initially extracted set of 13 common roots contains one zero which in the general unrestricted dynamic model (4), shortens the lag length by one for every variable. This was confirmed in an F-test on the coefficients of the variables lagged 13 months, which gave a value of 0.356 for the $F(3,210)$ test-statistic which is not significant. Reestimating equation (4) allowing this time the scalar polynomials to be of order 12 and testing for a common factor polynomial of degree m we obtained the following values for the Wald criteria.

Wald Criteria

m	Wald Criteria	Degrees of Freedom	Differences	Degrees of Freedom
11	1.456	22	-	-
12	8.960	24	7.504*	2

* Significant at the 5% level

As can be seen from the above table, the application of the "common factor" analysis in the simplified version of (4), gives similar results to the ones previously obtained.

The above empirical analysis indicates that the overall dynamics in the export demand equation for all goods is of 12th order. The model selection procedure which starts from the static form of the basic equation leads to a model where the overall dynamics in the equation is expressed only through systematic dynamics (model (3)). On the contrary, starting from the general unrestricted dynamic model, the common factor analysis based on the Wald criteria suggests that the overall dynamics in the equation is expressed only through error dynamics (model (2)).

Turning now to a test of the seasonal adjustment of the data, we reestimated equation (3) using unadjusted data. The F-test constructed from the corresponding residual sums of squares gave a value of 4.151 for the F(11,246) test-statistic, which is highly significant. Finally, according to the likelihood ratio test performed in the above model, the log-linear form is preferred to the linear form.

The long-run income elasticity obtained from model (3) is 2.70 showing that with respect to the index of industrial of the O.E.C.D. countries, Greek exports are income elastic. On the other hand, the long-run price elasticity is -1.325, and while being of reasonable magnitude (in view of the price elasticites of the individual categories), it is nevertheless insignificant.

It appears then, that in the aggregate export demand equation the relative prices are an unimportant factor and add almost nothing to the explanation of exports. Thus, although on the basis of the individual export demand equations prices are quite important in determining changes in the volume of exported commodities, the aforementioned failure of the aggregate function to reveal a price effect, shows the value of disaggregation.

To derive the aggregate export demand elasticities from the disaggregated demand functions, we follow again the method used by Barker (1970), already described in the previous chapter in relation to the aggregate import demand function. Since in most of our disaggregated export demand functions, the preferred model is of the static form, it is only this form of export demand equation that is considered. In particular, for the categories, food, beverages and tobacco, raw materials and manufactures, the elasticities used for the calculation of the aggregate export demand elasticities, result from the corresponding preferred original equations, autoregressive error specifications. The export demand elasticities for chemicals are obtained from the preferred autoregressive form which was reestimated using as activity variable the index of industrial production of the O.E.C.D. countries, instead of the index of chemical production which was initially employed.

The table below shows the derived aggregate elasticities which were obtained using the average and the 1976 shares of each export category in the total exports. From this table it can be seen that the derived aggregate activity elasticity is in the region of 1.7 (using the average shares) which is quite close to the direct estimate of 1.733 from model (2). But, the derived aggregate price elasticity is very low compared with the individual price elasticities. This is due to the fact that all export categories, apart from beverages and tobacco, have very low export price distribution elasticities. If we ignore the price distribution elasticities, the derived aggregate price elasticity, using the average and the 1976 shares, is -0.56 and -0.48 respectively.

However, the elasticities obtained from the disaggregated equations are downward-biased because these do not allow for exports of the omitted from the analysis SITC categories which are included in total exports (see section 1). These "excluded" exports, which were less than 5 percent of total exports during the period 1954 - 1971, accounted for about 21 percent during the period 1972 - 1976. In particular, exports of fuels and lubricants (SITC: Section 3) which were not exported at all before 1972, averaged to 8.6% in the value of the total exports of Greece during the period 1972 - 76 (the main

TABLE 1

INCOME AND PRICE ELASTICITIES FOR FIVE CATEGORIES OF EXPORTS AND TOTAL EXPORTS

Export Category	Short - run Income Elasticity	Short - run Price Elasticity	Shares in Total Exports		Estimated Relative Price Distribution Elasticity
			Average 1954-76	1976	
Food	1.763	-0.532	0.237	0.206	0.195 (0.082)
Beverages and Tobacco	0.646	-0.355	0.188	0.092	1.134 (0.161)
Raw Materials	0.923	-1.126	0.145	0.087	0.253 (0.074)
Chemicals	1.818	-1.200	0.051	0.041	0.236 (0.155)
Manufactures	3.918	-0.586	0.238	0.325	0.313 (0.104)
Sum of individual shares			0.859	0.751	
Total estimate derived using					
(i) Average shares	1.698	-0.200			
(ii) 1976 shares	1.851	-0.154			

exported item of this group is petroleum products refined). Exports of machinery and transport equipment (SITC : Section 7), which accounted for about 1% of total exports before 1972, they averaged to 3.1% in the years 1972 - 76 (this group mainly includes equipment for distributing electricity, and motor vehicles for the transport of goods or materials and special purpose motor vehicles). Finally, exports of miscellaneous manufactured articles (SITC: Section 8) which accounted for about 1% in total exports during the years 1954 - 71, they averaged to 9.3% in the period 1972 - 76 (this group mainly includes outer and under garments, and footwear).

In view of the above, as can be seen from the above table, the groups of exported commodities previously examined, account for an average of 85.9% of total exports during the whole sample period, while for 1976 they account only for 75.1%. Adjusting the shares to add up to unity the derived activity elasticity (or the aggregate activity elasticity of the listed only groups), using the average and the 1976 shares, is 1.977 and 2.465 respectively. The derived short-run relative price elasticity becomes -0.233 and -0.205 respectively, and if we ignore the price distribution elasticities, the weighted price elasticity is -0.648 and -0.639.

3. <u>Summary of Findings</u>

A summary of the estimated elasticities of the export demand equations for the major groups of exported commodities and the total exports are presented in the table below. Long-run activity and price elasticities are presented for those groups of commodities for which dynamic export demand functions were estimated.

The results obtained from the export demand equations of Greece may be summarized as follows.

(i) Exports of food to O.E.C.D. countries, which provide most of the demand for such goods, are elastic with respect to the industrial production of these countries. Taking also into consideration that industrial production of the O.E.C.D. countries has grown faster than their income (during the sample period total industrial production and national income increased at an average annual rate of 4.8% and 4.1% respectively) the income elasticity of exported food would be placed in the neighborhood of 2.

(ii) Tobacco exports (which account for more than 85% of exports of beverages and tobacco) have grown slower than the production of food, beverages and tobacco of the importing countries. However, since Greece exports only cigarette leaf tobacco, it is expected that the activity elasticity of tobacco exports with respect to the cigarette production of the O.E.C.D. countries, should be around one.

(iii) The volume of exported raw materials has grown as fast as the level of industrial production of the O.E.C.D. countries. However, with respect to the income of the O.E.C.D.

TABLE 2

EXPORT DEMAND ELASTICITIES FOR GOODS
(Monthly Data 1954-1976)

Exports of:	Industrial Production Elasticity		Relative Price Elasticity	
	Short-run	Long-run	Short-run	Long-run
Food	1.763	1.888	-0.532	-0.755[a]
Beverages and Tobacco	0.646 0.829[b]		-0.355	
Raw Materials	0.923	0.955	-1.126	-1.334
Chemicals	1.360[c] 1.818		-1.219	
Manufactures	3.918	4.208	-0.586	-1.677
All Goods	1.733	2.700		-1.325[d]

[a] Significant at the 8% level

[b] With respect to industrial production of food, beverages and tobacco of OECD

[c] With respect to chemical production of OECD

[d] Not significant at the 10% level

countries, the export demand elasticity for raw materials appears to be greater than one.

(iv) Regarding Greece's exports of chemicals, it appears that the foreign demand elasticity for these products with respect to the chemical production of the importing countries, is somewhat above unity. On the other hand, the export demand elasticity for manufactures, with respect to the index of industrial production of the O.E.C.D. countries is around 4. This high activity elasticity for manufactures, is partly attributed to the rapid growth of the Greek manufacturing industries resulting in an increasing supply of the commodities in question.

(v) Finally, taking the relative price elasticities of the individual groups of commodities, at their face value, their magnitudes on the whole indicate that Greek exports face noticeable price competition in foreign markets.

CHAPTER VII

CONSIDERATIONS OF A JOINT ESTIMATION OF THE IMPORT AND EXPORT DEMAND EQUATIONS

Up to this point we have estimated each import and export demand equation separately using least squares (or non-linear least squares according to the error specification), taking no account of any correlation across the equations that might exist (see chapter IV, section 2.2.).

In matrix form, a set of the above estimated n import (or export) demand equations can be written as follows:

$$(1) \qquad \mathbf{y}_t = \sum_{i=1}^{m} \mathbf{B}_i \mathbf{y}_{t-i} + \mathbf{C}\,\mathbf{x}_t + \mathbf{u}_t \qquad (t \quad 1, \ldots, T)$$

where

$$\mathbf{u}_t = \sum_{i=1}^{\ell} \mathbf{R}_i \mathbf{u}_{t-i} + \boldsymbol{\varepsilon}_t$$

and $\boldsymbol{\varepsilon}_t$ is NID $(\mathbf{0},\boldsymbol{\Sigma})$. $\mathbf{C}$ and the (diagonal) $\mathbf{B}_i$ are matrices of coefficients, and $\mathbf{y}_t$ and $\mathbf{x}_t$ are n×1 and p×1 vectors of observations on the dependent variables and all explanatory variables, which are included in the import (or export) demand equations, at time t.

The application of least-squares to (1) requires that the matrices $\mathbf{R}_i$ and $\boldsymbol{\Sigma}$ are diagonal and this is the approach followed in the previous chapters. However, there may be a contemporaneous correlation between the disturbances reflecting the omission of any common factors that affect the behaviour of imports (or exports) of all or some commodity groups (e.g. foreign exhange reserves of the country for imports and world excess demand for exports). Whenever we have such a situation the matrix $\underset{\sim}{\Sigma}$ is non-diagonal and the equations are only seemingly unrelated. In this case, more efficient estimates will be obtained if the equations in (1) are estimated jointly using a joint generalized least squares type of estimation (see Zellner (1962 and 1963)). Another possibility is the existence of non-zero correlations between lagged disturbances from a pair of equations. This type of correlation would arise if the effect of the omitted variables that are common to both equations, take different times to be demonstrated from one equation to the other. In this case an extension of Zellner's joint estimation technique can be applied (see Park (1967)).

The above situations can be detected in practice by the use of residual cross correlation functions. That is, for a pair of equations, we compute

$$\rho_{12}(k) = \frac{\text{cov}\ (\varepsilon_{1,t}\ ,\varepsilon_{2,t-k})}{\sqrt{\text{var}(\varepsilon_1)\cdot\text{var}(\varepsilon_2)}}$$

for

$$k = \ldots\ , -1,\ 0,\ 1,\ \ldots$$

where ε_i denotes the (least-squares) residuals in equation i. We then plot $\rho_{ij}(k)$ against k, with a band of $\pm 2/\sqrt{T}$ (the asymptotic standard error of a correlation coefficient between two independent series being $1/\sqrt{T}$). Thus a departure from the null hypothesis of no correlation is indicated whenever $\rho_{ij}(k)$ exceeds this critical value.

Such residual cross correlation functions are shown below for different pairs of import and export demand equations. For convenience we shall present first the cross correlograms of the import demand functions and then the cross correlograms of the export demand functions.

1. Residual Cross Correlation Functions of import Demand Equations

The residuals used for the construction of the cross correlograms are obtained from the preferred forms already described in chapter V. Since two different model selection procedures have been applied for each import (or export) category, whenever there are more than one preferred form , the residuals have been obtained from that equation which in terms of residual variance provides a better fit.

To preserve comparability in the residuals, all equations have been estimated from seasonally adjusted (by the least-squares method) logarithms of the variables. Thus, for imports of fuels, where the estimation of general autoregressive forms was virtually impossible because of convergence difficulties when we used logarithms of the variables, the residuals were obtained from the preferred equation which was selected on the basis of simple autoregressive error processes (see chapter V, section 2.3.).

For imports of machinery and transport equipment the estimation of general autoregressive error forms was impossible only for processes of order greater than 9. So, among the autoregressive forms which were possible to estimate, an autoregressive error specification of general 8th order form was preferred according to the likelihood ratio tests[1]. Subsequently, the residuals were obtained from that equation.

In view of the discussion in chapter IV, section 1.4., the residual cross correlation functions $\rho_{ij}(k)$ have been estimated for k = -13, ... , -1, 0, 1, ... , 13. The subscripts i, j refer to the SITC categories and therefore take the values 0, 2, 3, 5, 6, and 7 (we recall that these SITC sections correspond to the following groups of imported commodities: food, raw materials, fuels, chemicals, manufactures, and machinery and transport equipment). Cross correlograms are shown only for these pairs of import demand equations for which the correlation coefficient between disturbances was

---

[1] The estimated equation is:

$$\log(M_{mt}) = \underset{(0.957)}{3.384} - \underset{(0.145)}{0.785} \log(P^m_{mt}/P^d_g) + \underset{(0.135)}{1.011} \log(IP)$$

$$u_t = \underset{(0.062)}{0.272}\, u_{t-1} + \underset{(0.064)}{0.024}\, u_{t-2} + \underset{(0.064)}{0.232}\, u_{t-3} +$$

$$+ \underset{(0.065)}{0.064}\, u_{t-4} - \underset{(0.066)}{0.049}\, u_{t-5} + \underset{(0.064)}{0.100}\, u_{t-6} +$$

$$+ \underset{(0.066)}{0.049}\, u_{t-7} + \underset{(0.063}{0.174}\, u_{t-8} + \varepsilon_t$$

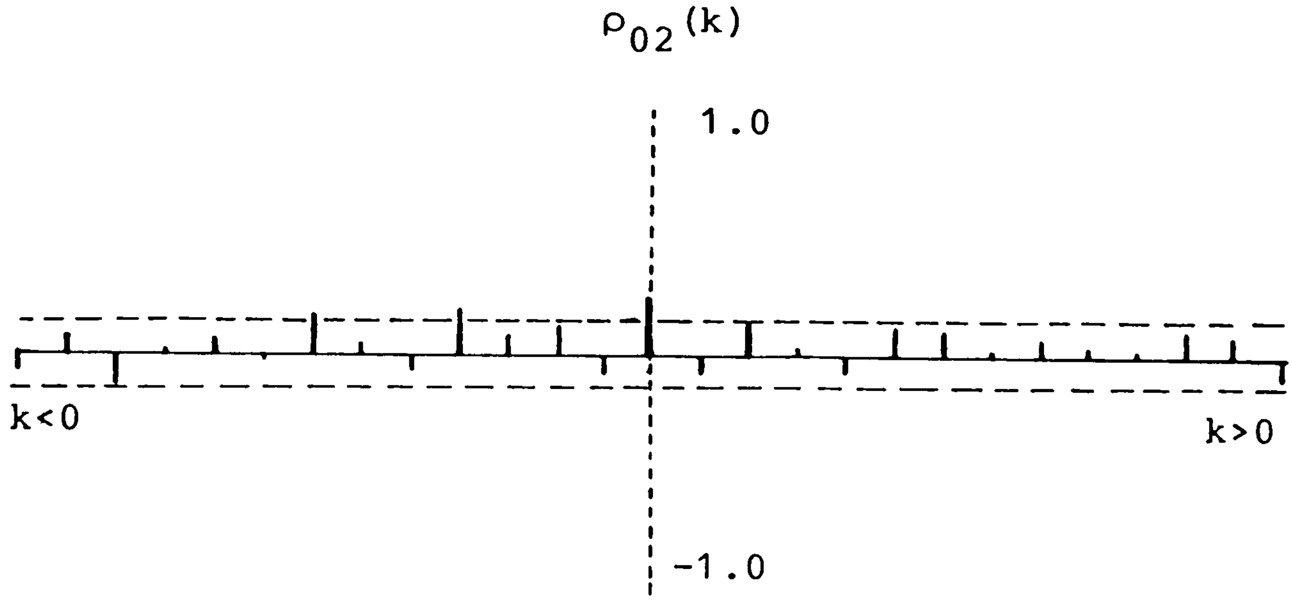
$$n = 263,\ R.S.S. = 5.343,\ X^2(15) = 3.342$$

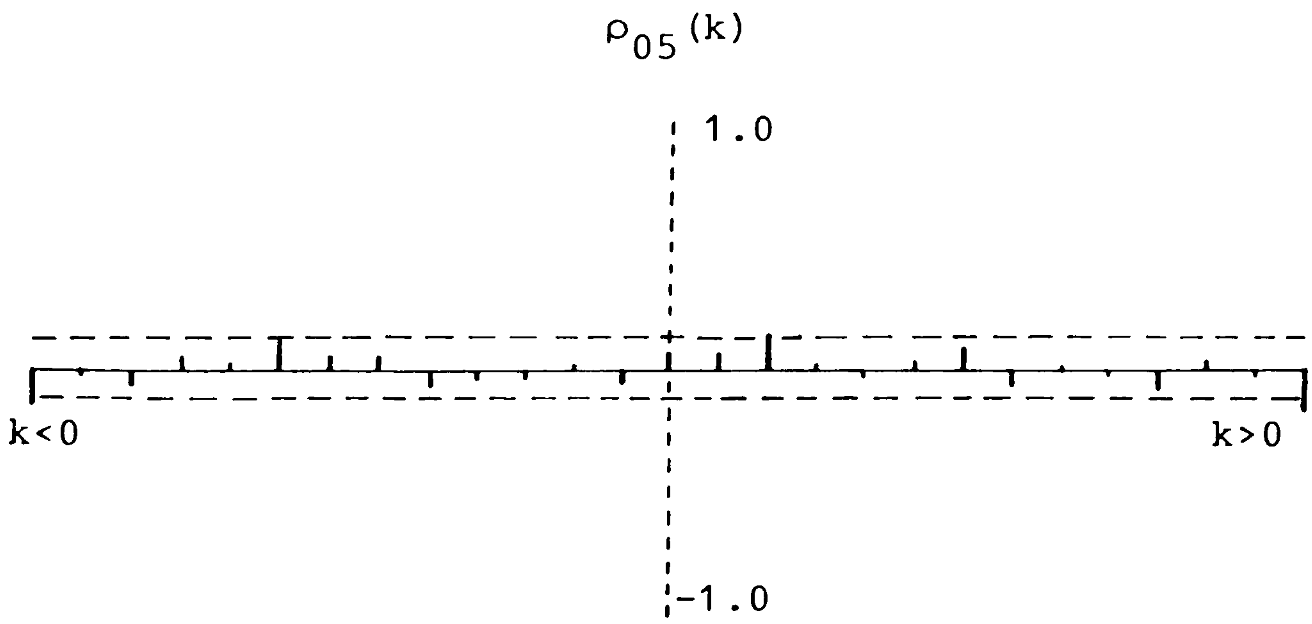
which was accepted against the unrestricted transformed equation ($X^2(16) = 17.981$).

shown to be significant for at least one value of k. The residual cross correlation functions now follow with a band of $\pm$ 0.1265 ($= \pm\ 2/\sqrt{250}$).

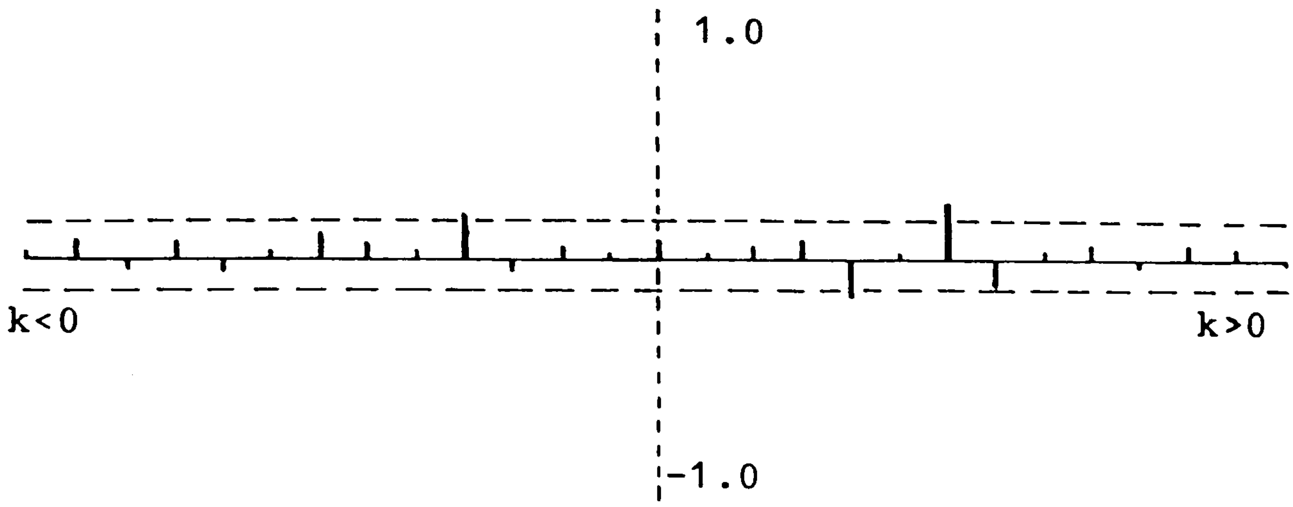
FIGURE 1

RESIDUAL CROSS CORRELATION FUNCTIONS OF IMPORT DEMAND EQUATIONS

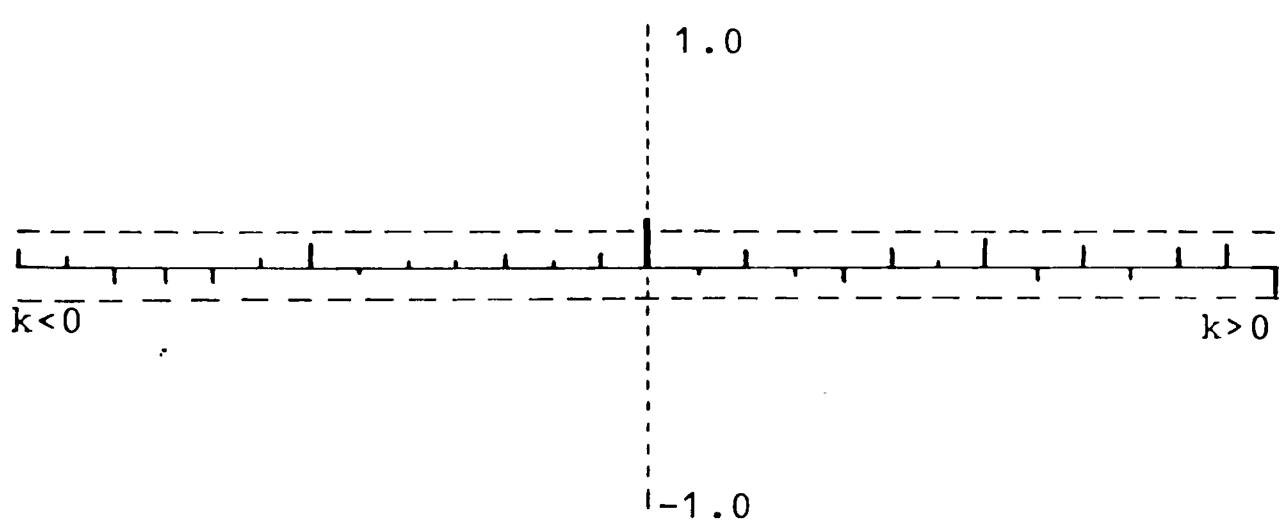




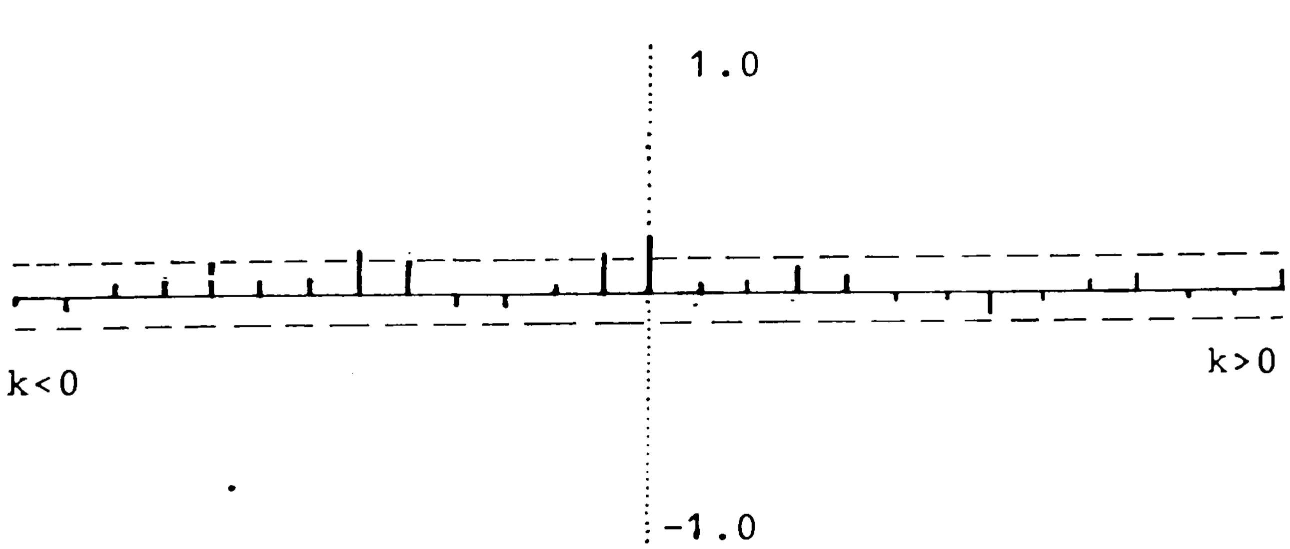
$\rho_{06}(k)$

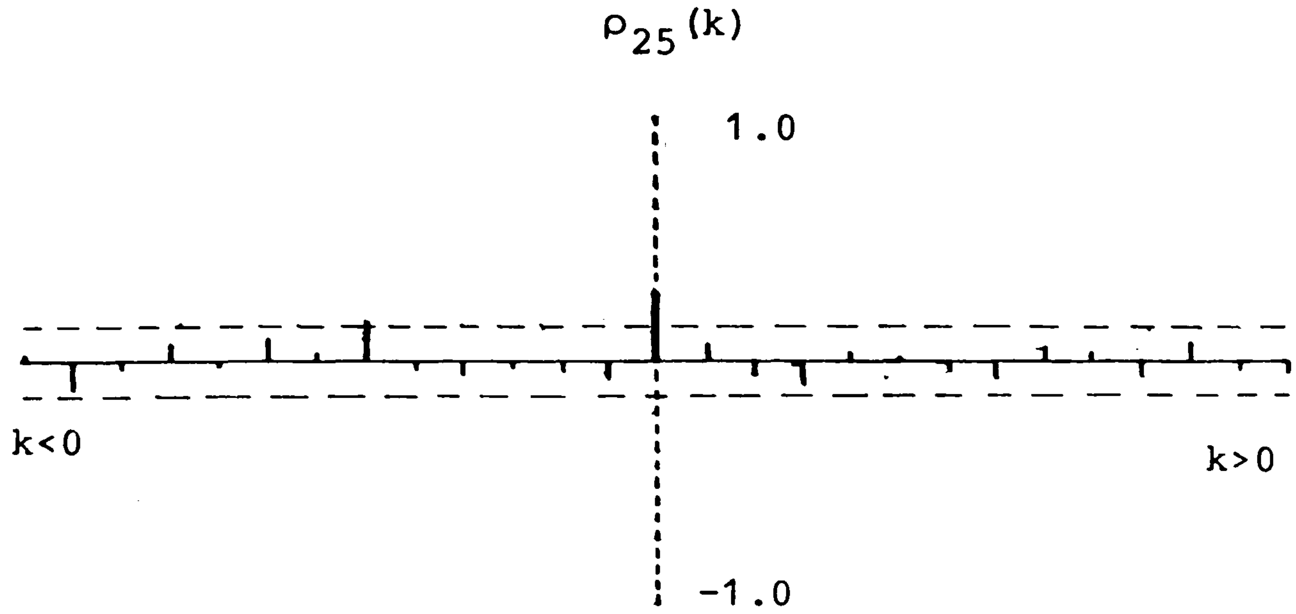


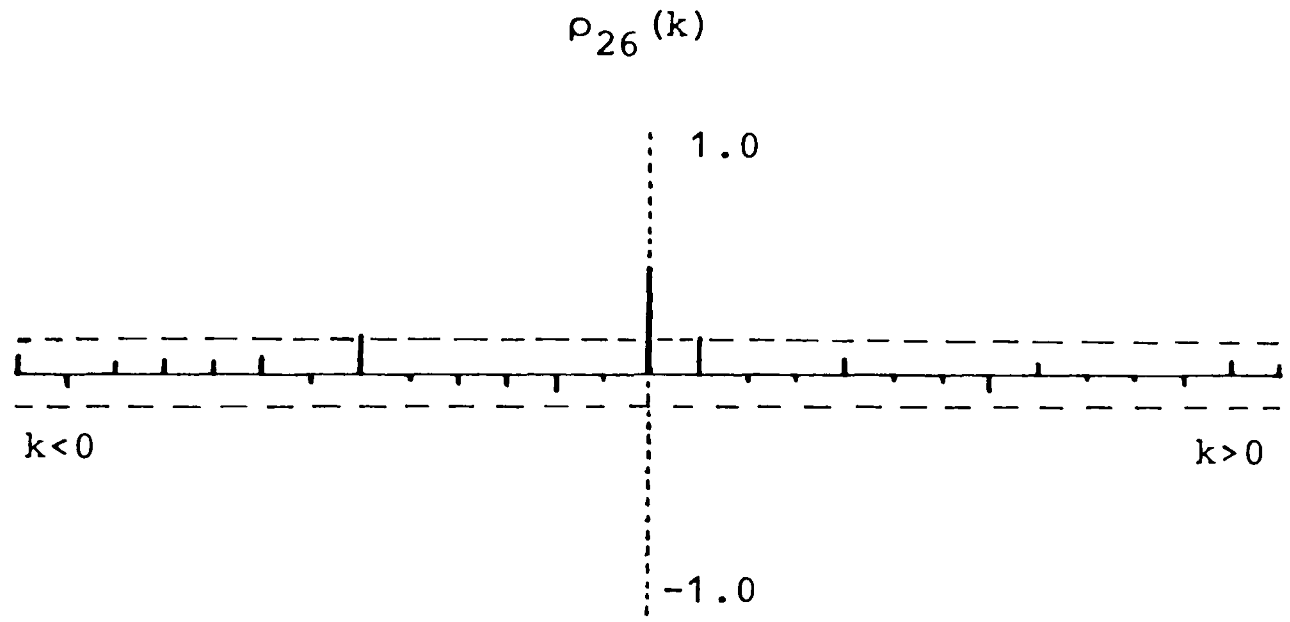
$\rho_{07}(k)$

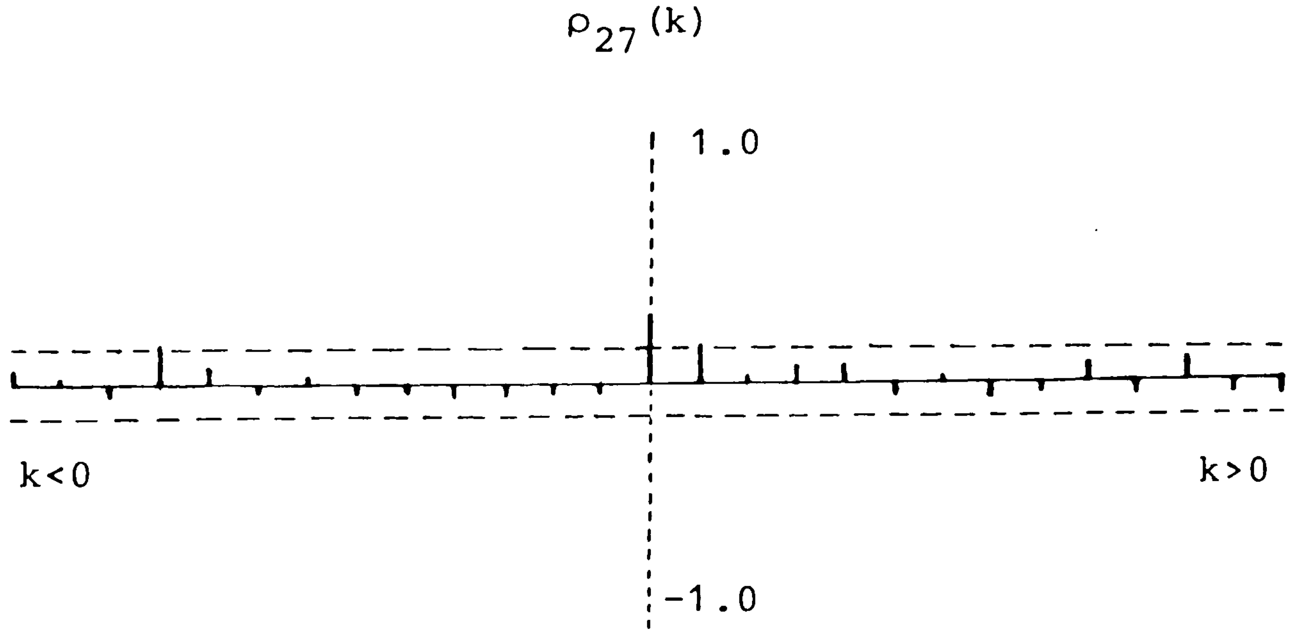


$\rho_{23}(k)$

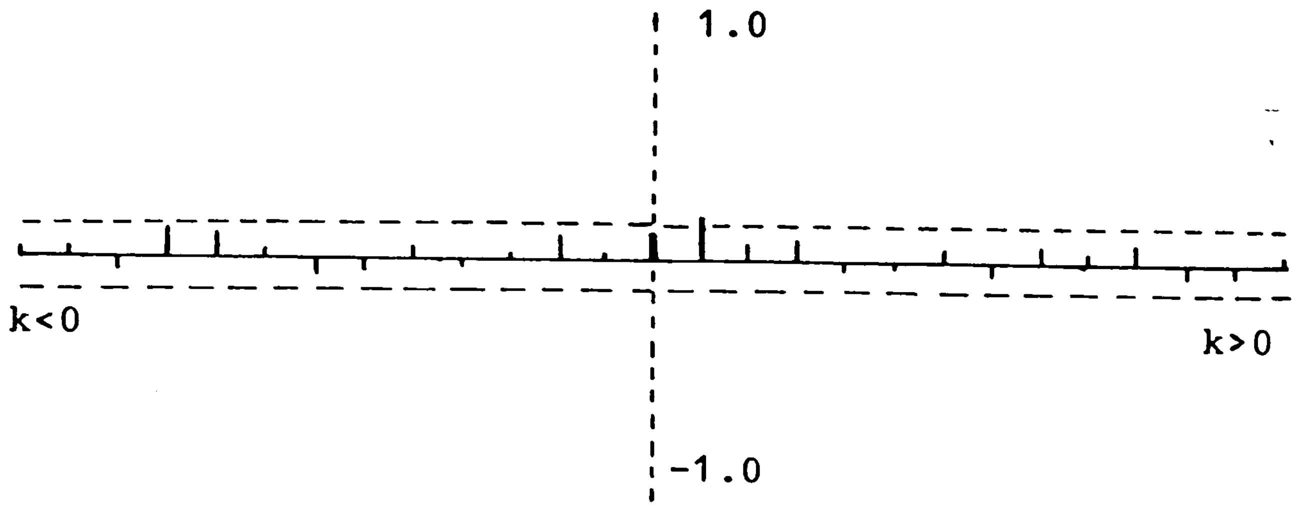



$\rho_{25}(k)$
1.0
k<0
k>0
-1.0

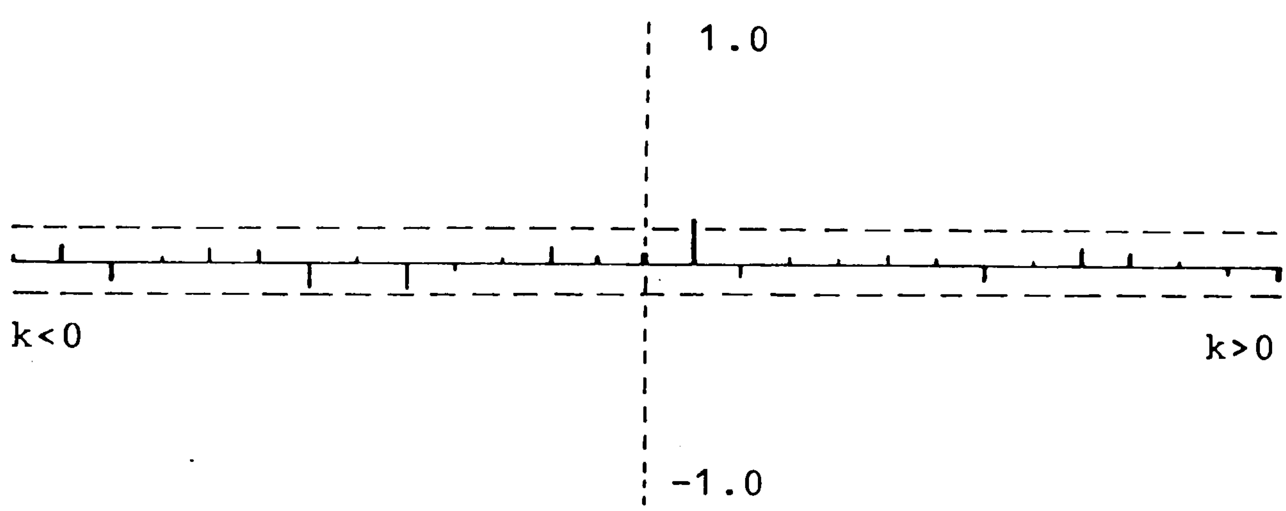

$\rho_{26}(k)$
1.0
k<0
k>0
-1.0


$\rho_{27}(k)$
1.0
k<0
k>0
-1.0

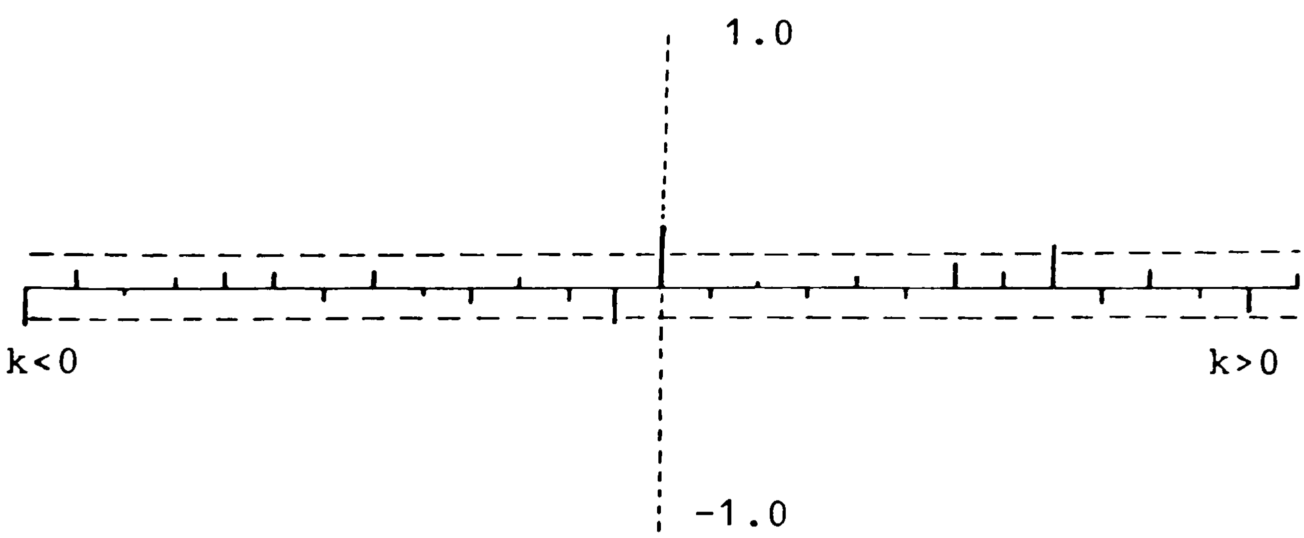
$\rho_{36}(k)$

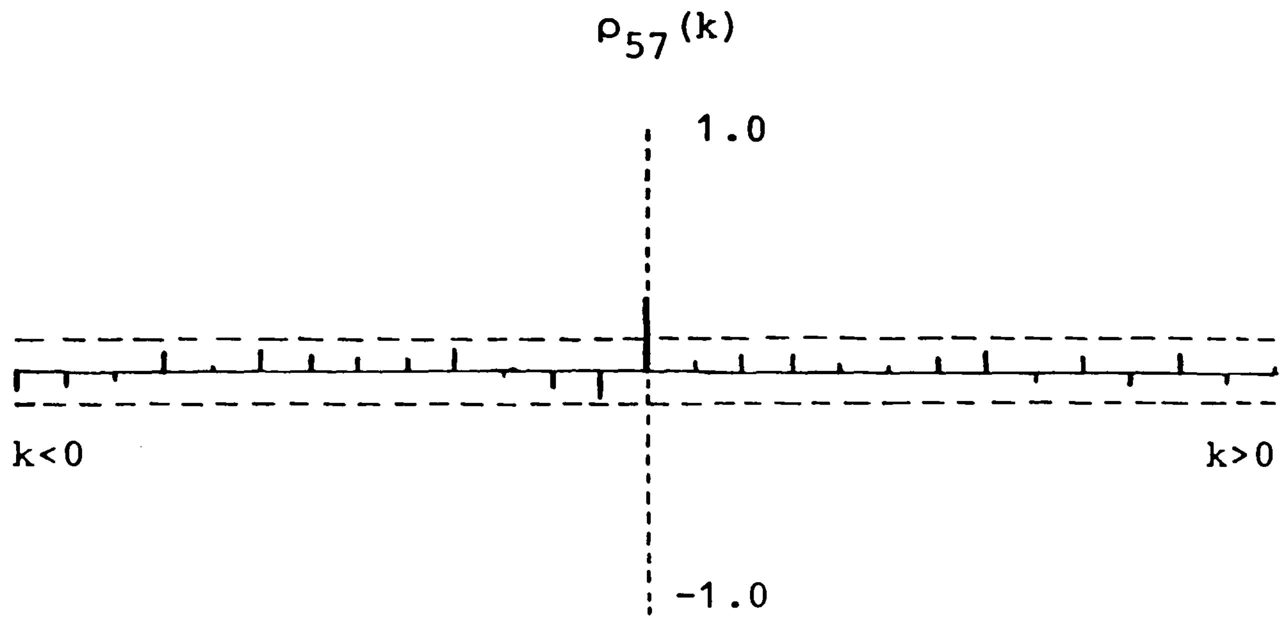


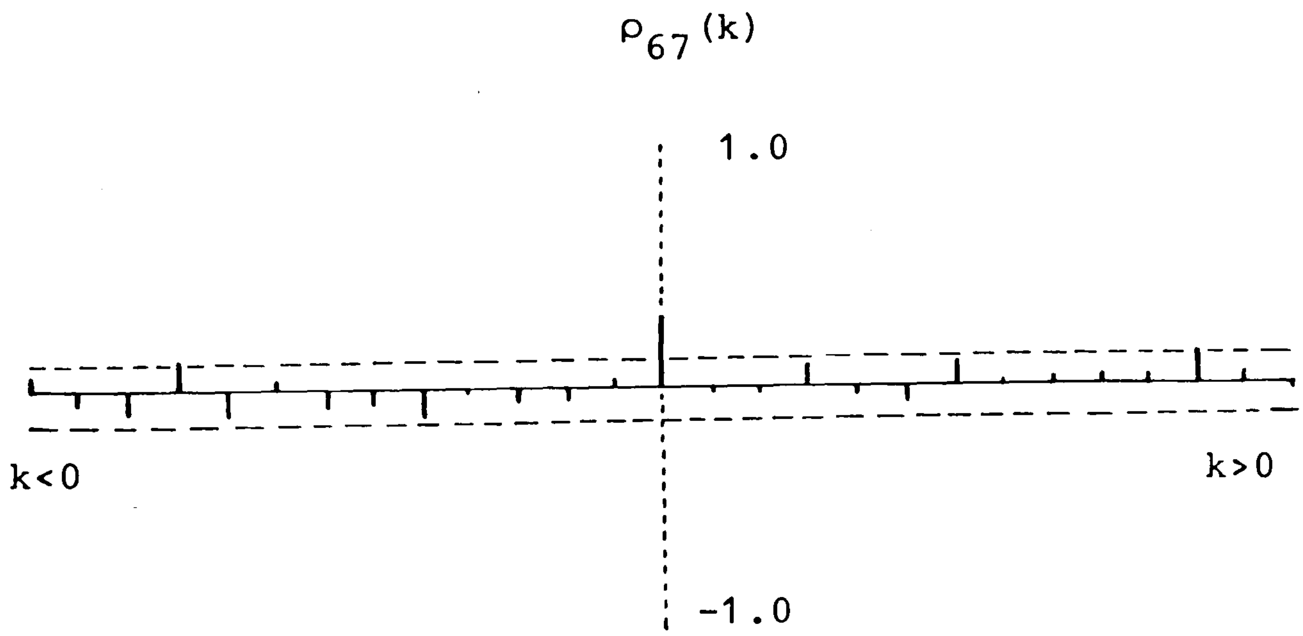
$\rho_{37}(k)$



$\rho_{56}(k)$




ρ57(k)
1.0
k<0
k>0
-1.0


ρ67(k)
1.0
k<0
k>0
-1.0

A number of interesting points arise from the above residual cross correlation functions. At least one $\rho_{ij}(k)$ was found to be greater than twice its standard error, in thirteen cases out of 15 (= $\binom{6}{2}$) possible pairs of equations. A significant contemporaneous correlation between the disturbances was found in 9 cross correlograms. In all these 9 cases the correlation coefficient $\rho_{ij}(k)$ takes the largest value for k = 0. Also, in 2 out of these 9 pairs of equations only $\rho_{ij}(0)$ was significant.

The largest contemporaneous correlations between disturbances were found in the following pairs of equations: (2,5), (2,6), (2,7), i.e. between imports of raw materials and respectively, imports of chemicals, manufactures, and machinery and transport equipment. These contemporaneous correlations are not unreasonable if we take into consideration that the industrial sector is the main source of demand for these groups of commodities (see the relevant sections in chapter V for a description of the composition of these commodity groups). As already mentioned in chapter V, due to data limitations changes in stocks and private fixed investment have not been introduced in the import demand equations for the above commodities. Therefore, these "omitted" variables may be the cause of the contemporaneous correlations across the above equations. Similarly, we found significant contemporaneous correlations between the disturbances in the pairs of equations (5,6), (5,7) and (6,7).

With regard to the correlations between disturbances in the pairs of equations (0,2), (0,5), (0,6), and (0,7), i.e. between imports of food and all other groups of imported commodities (except fuels), although significant they hardly justify the application of a joint estimation technique of these pairs of equations. The significant values of $\rho_{0j}(k)$, $j = 2, 5, 6$, and 7, probably reflect the omission of some general common factor affecting the imports of these groups of commodities, as for example the ability of the country to import. As mentioned earlier in chapter V, section 1, the ability of the country to import was measured by the receipts of the country from exports of goods and services. However, this variable was not introduced into the import demand equations because due to the high intercorrelation between that variable and the index of industrial production, it failed to produce significant results.

Finally, the lower correlations between disturbances appear in the pairs of equations (2,3), (3,6) and (3,7), i.e. between imports of fuels and imports of raw materials, manufactures, and machinery and transport equipment. Among these three pairs of equations the highest correlation appears in the pair (2,3), i.e. between imports of raw materials and imports of fuels, and is contemporaneous.

From the above we conclude that in general any correlation between the disturbances of the import demand functions tends to be contemporaneous rather than lagged. As can be seen from the table below, among these contemporaneous correlations, the largest value is 0.425 occuring in the pair of equations (2,6), i.e. between imports of raw materials and imports of manufactures, while the rest (significant ones) take values within the range 0.166 to 0.301 .

TABLE 1

CONTEMPORANEOUS CORRELATIONS BETWEEN THE RESIDUALS OF THE IMPORT DEMAND EQUATIONS

	SITC Categories					
	0	2	3	5	6	7
0	1.000	0.219	0.010	0.062	0.087	0.166
2		1.000	0.204	0.289	0.425	0.285
3			1.000	0.100	0.104	0.029
5				1.000	0.247	0.301
6					1.000	0.299
7						1.000

The question now arising is whether the observed correlations between the disturbances, mostly contemporaneous, are high enough in order to indicate the application of a joint generalized least squares type of estimation. It can be shown, that if the matrix of explanatory variables is the same for each equation, then the joint generalized least squares estimators are reduced to ordinary least-squares estimators. Thus, the gain in efficiency yielded by the Zellner estimator over the least-squares estimator increases directly with the correlation between the disturbances from the different equations and inversely with the correlation between the different sets of explanatory variables (Johnston (1972), p p. 240 - 241).

In our import demand equations, imports are determined in terms of activity variable and relative prices. The activity variable entering all our import demand functions, is the index of total industrial production of Greece. The index of chemical production has been employed as activity variable, instead of the index of industrial production, only in the case of imports of chemicals. However, these two indices are highly correlated (r = 0.991). Taking now the relative prices, it appears from the table below, that there is a noticeable correlation between the relative prices of the various groups of imported commodities (RP_i stands for the relative prices of the ith SITC section).

TABLE 2

CORRELATION COEFFICIENTS BETWEEN RELATIVE IMPORT PRICES

	RP_0	RP_2	RP_3	RP_5	RP_6	RP_7
RP_0	1.000	0.327	0.679	0.645	0.702	0.346
RP_2		1.000	0.120	0.656	0.120	0.803
RP_3			1.000	0.425	0.601	0.002
RP_5				1.000	0.522	0.588
RP_6					1.000	0.100
RP_7						1.000

Also the lagged values of the dependent variables enter our import demand equations. However, lagged values of the dependent variables appear among the regressors in two only equations (import demand equation for raw materials and import demand equation for fuels).

In view of the above, there is not enough evidence to support the view that there will be any gain in efficiency if the selected forms of the import demand equations are jointly estimated.

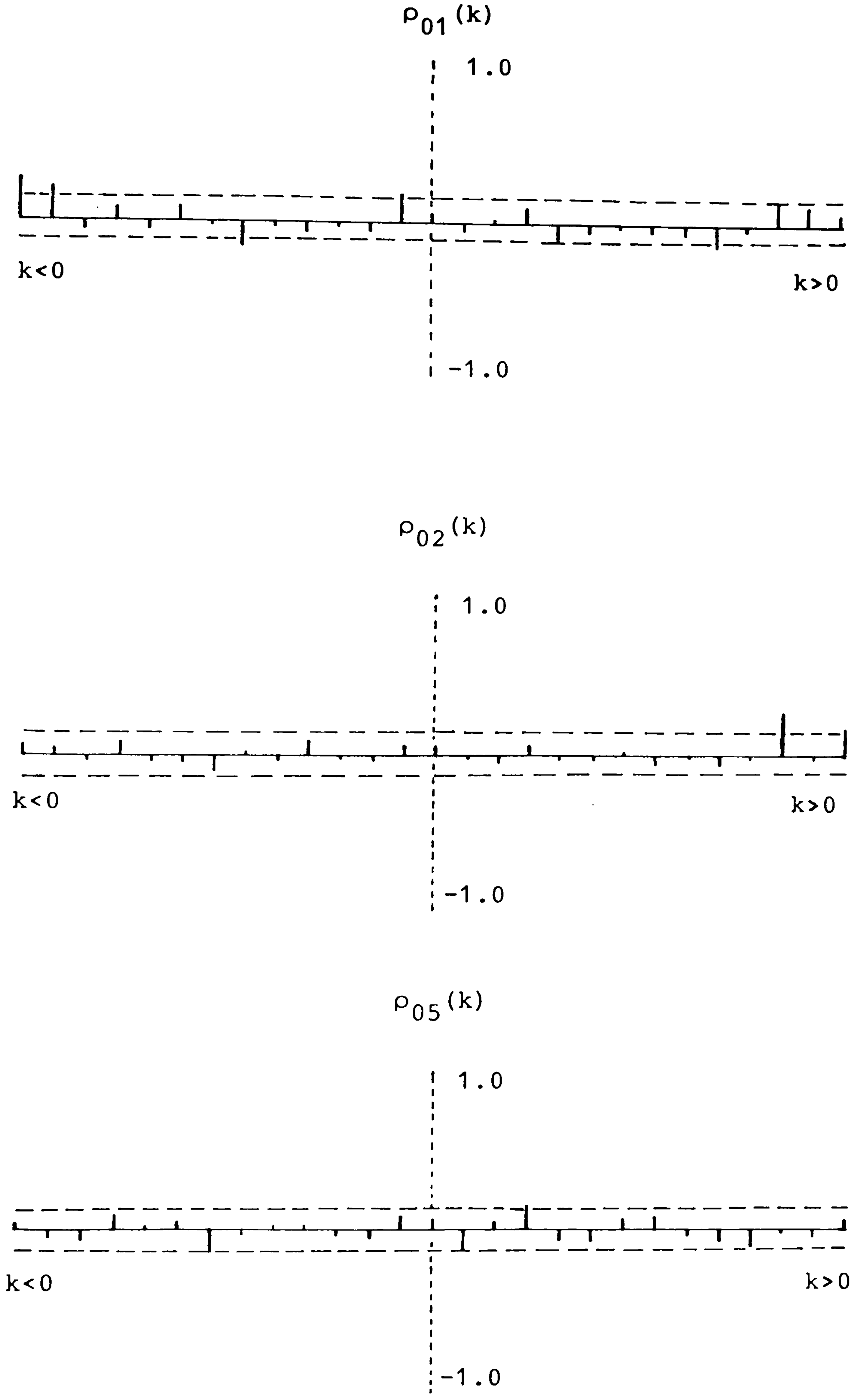
2. Residual Cross Correlation Functions of Export Demand Equations

The residuals to be used for the construction of the cross correlograms, have been obtained from the preferred forms of the export demand equations, already described in chapter VI. Here again, all equations have been estimated from seasonally adjusted (by the least-squares method) logarithms of the data. Cross correlograms are shown only for these pairs of export demand equations for which the correlation coefficient between disturbances $\rho_{ij}(k)$, was shown to be significant for at least one value of k, k = -13, ..., -1, 0, 1, ..., 13. The subscripts i, j take the values 0, 1, 2, 5 and 6 which refer to the SITC groups of exported commodities, food, beverages and tobacco, raw materials, chemicals, and manufactures, respectively.

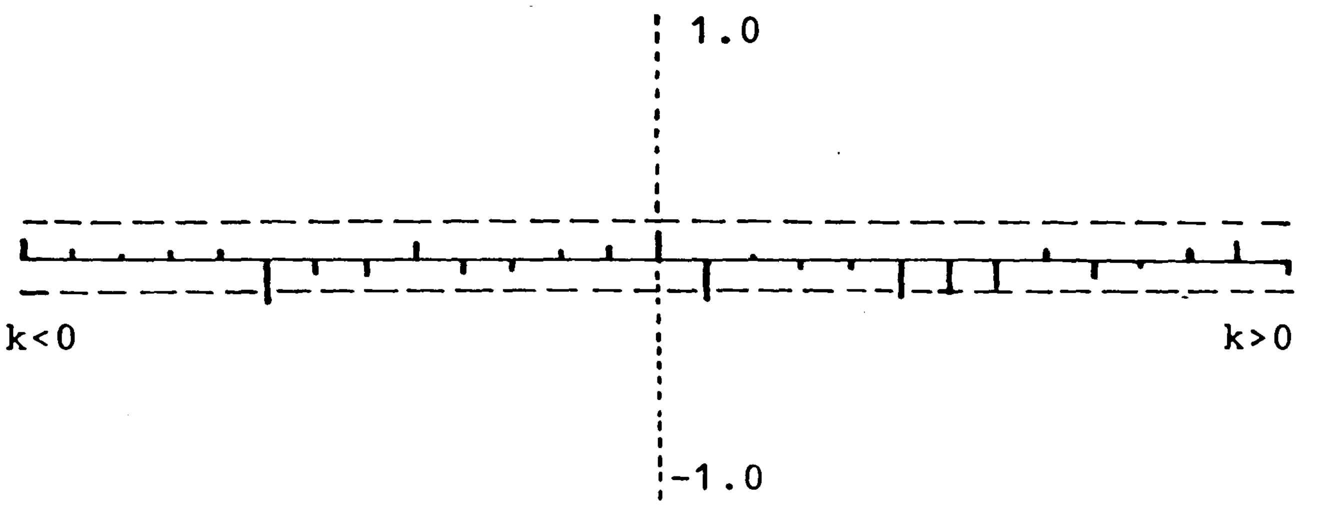
The residual cross correlation functions now follow with a band of $\pm 2/\sqrt{T}$ (= ±0.1265).

FIGURE 2

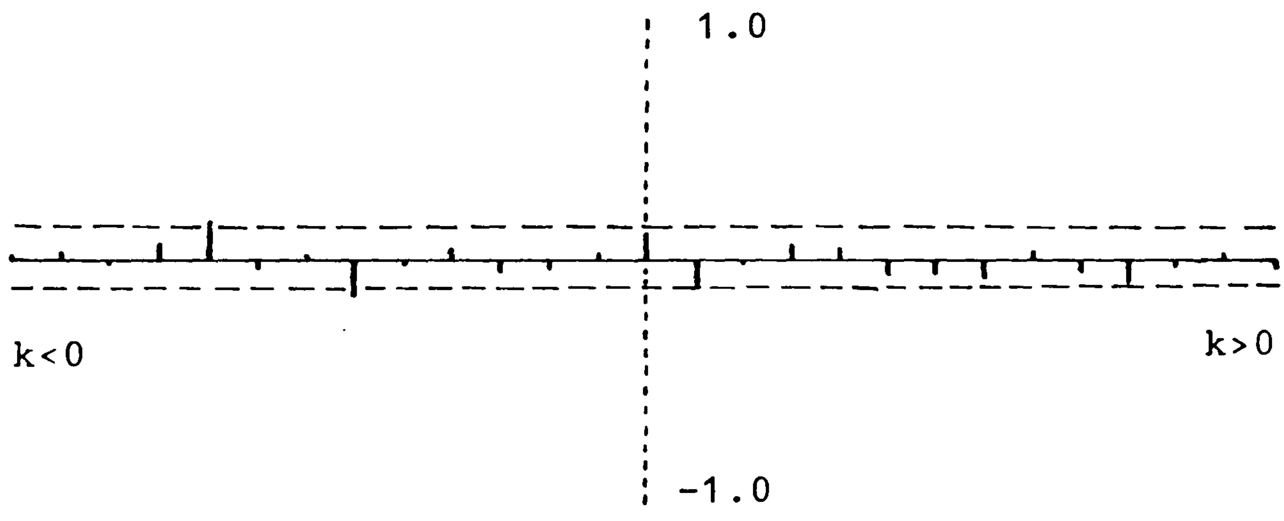
RESIDUAL CROSS CORRELATION FUNCTIONS OF
EXPORT DEMAND EQUATIONS



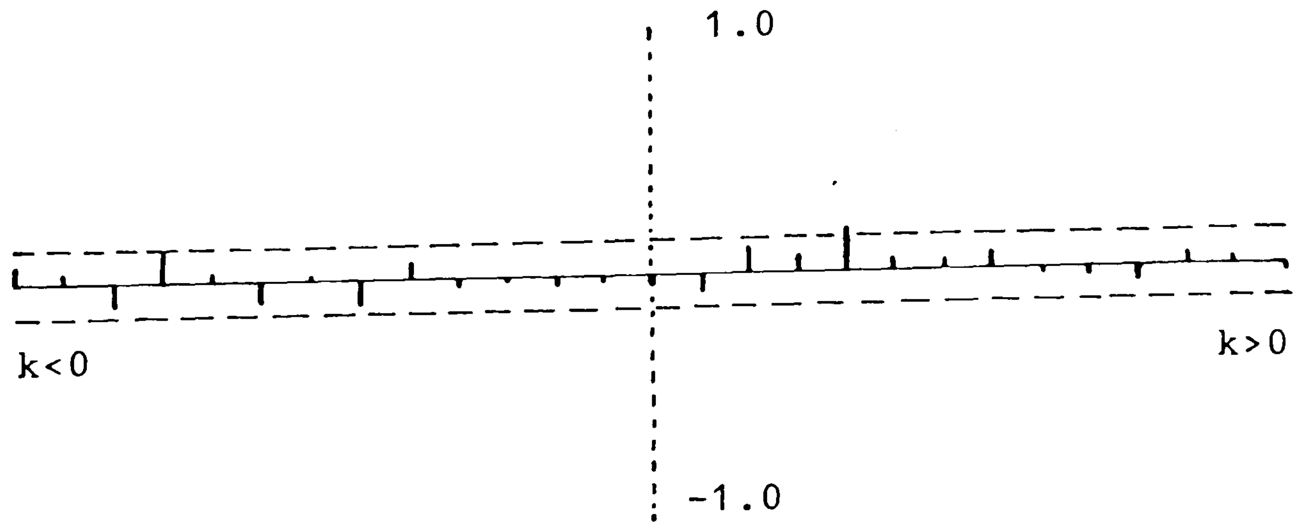
$\rho_{12}(k)$

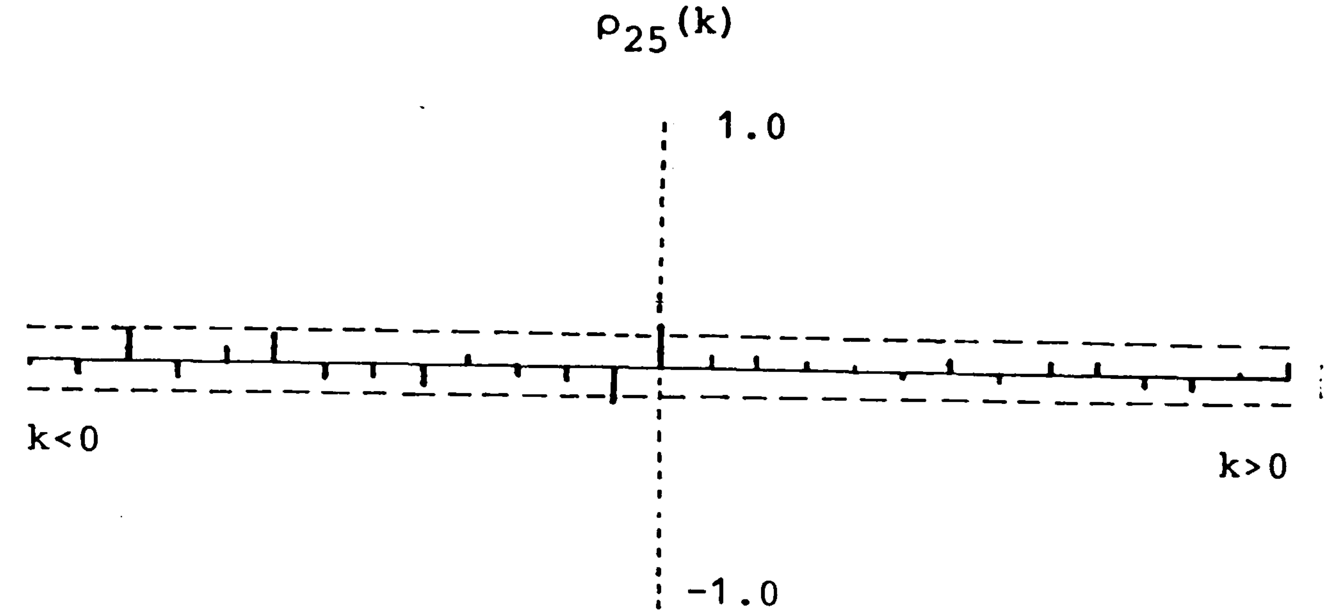


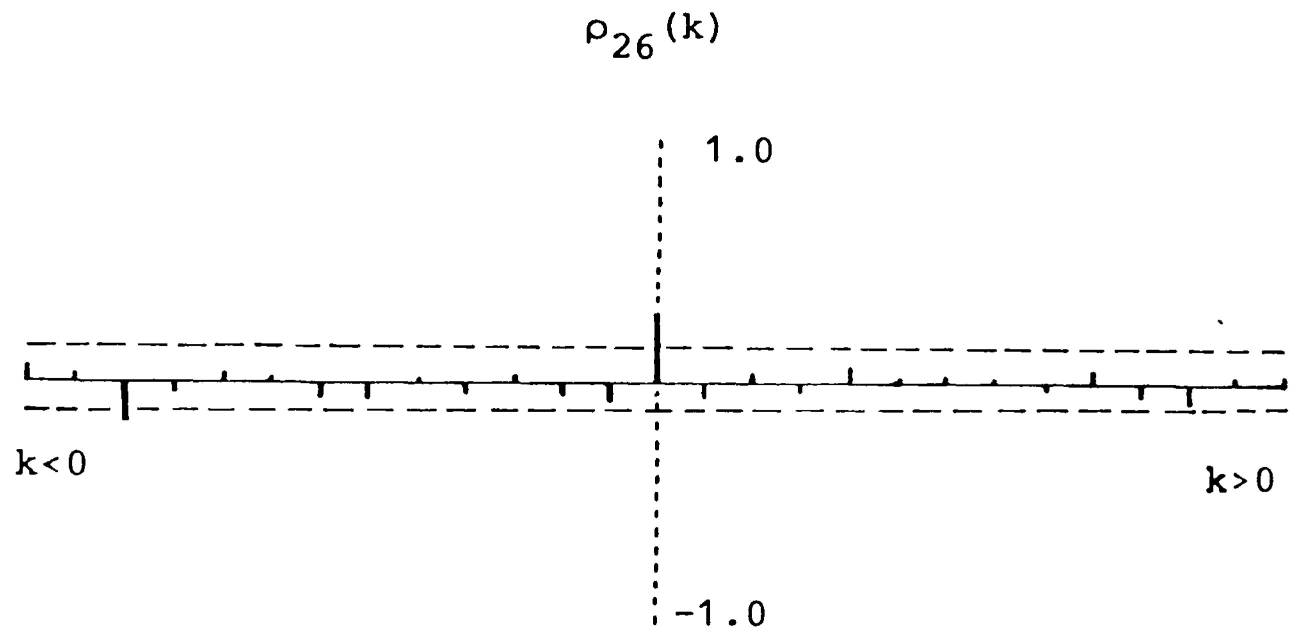
$\rho_{15}(k)$

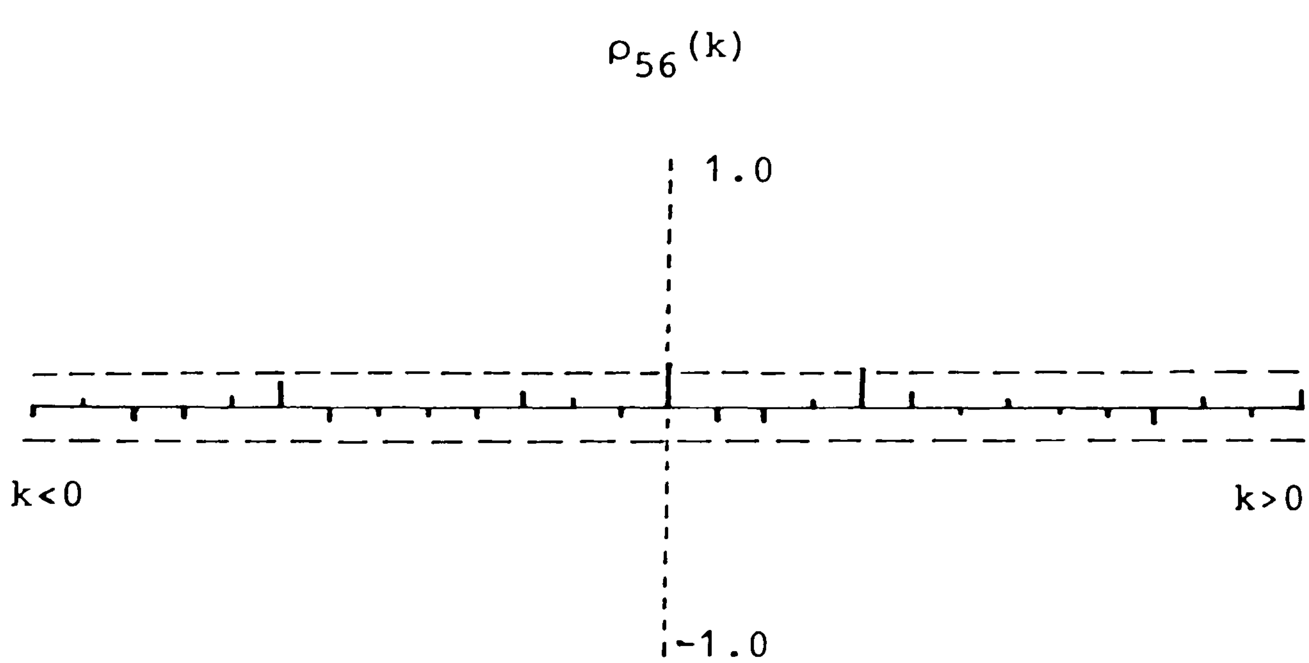


$\rho_{16}(k)$




$\rho_{25}(k)$
1.0
k<0
k>0
-1.0


$\rho_{26}(k)$
1.0
k<0
k>0
-1.0


$\rho_{56}(k)$
1.0
k<0
k>0
-1.0

The above residual cross correlation functions, give rise to the following points. At least one $\rho_{ij}(k)$ was found to be greater than twice its standard error, in nine cases out of 10 ($= \binom{5}{2}$) possible pairs of equations. A significant contemporaneous correlation between the disturbances was found in only 3 cross correlograms. In all these 3 cases the correlation coefficient $\rho_{ij}(k)$ takes the largest value for $k = 0$. Finally, in all nine residual cross correlation functions graphed above, at least one significant lagged correlation coefficient between disturbances, was found.

The largest lagged correlations between disturbances were found in the pairs of equations (0,1), (0,2), and (1,2), i.e. between exports of food, beverages and tobacco, and raw materials. These groups of commodities refer mainly to agricultural products (about 45% of exports of raw materials consist of raw cotton and hides and skins) and the significant values of $\rho_{ij}(k)$, $i,j = 0, 1, 2$, probably reflect the omission of some general common factor affecting the exports of these commodities. A lag effect lasting approximately one year can be observed from the pairs of export demand equations (0,1) and (0,2). In particular, for the pair of equations (0,1), it appears from the corresponding cross correlogram, that the residuals of equation 0 lead the residuals of equation 1 by approximately one year. On the contrary, for the pair of equations (0,2), we can see that the residuals of equation 2 lead those of equation 0 by the same period. However, the above observed leads do not seem to affect the lead/lag effects in the pair of equations (1,2) (the correlation coefficient $\rho_{12}(k)$, k = 23, 24, 25 and 26, was found insignificant).

Significant contemporaneous correlations between disturbances were found in the pairs of equations (2,5), (2,6) and (5,6), i.e. between exports of raw materials, chemicals and manufactures. These contemporaneous correlations are not unreasonable if we take into consideration that the industrial sector of the importing countries is the main source of demand for these groups of exported commodities (see the relevant sections in chapter VI for a description of the composition of these commodity groups).

Finally, the lower correlations between disturbances appear in the pairs of equations (0,5), (1,5) and (1,6), i.e. between exports of chemicals and exports of food and beverages and tobacco, and between exports of beverages and tobacco and exports of manufactures. These correlations between disturbances, although significant, are too small to justify the joint estimation of these pairs of equations, bearing also in mind the heterogeneity of these group pairs of exported commodities.

From the above we conclude that in general the observed correlations between the disturbances of the various export demand equations are quite small. From these contemporaneous and lagged correlations, the largest values amount to 0.291 and 0.276 and occur in the pairs of equations (2,6) and (0,2), i.e. between exports of raw materials and exports of manufactures and exports of food, respectively. The rest (significant) correlations between disturbances take values within the range 0.127 to 0.223.

With regard to the correlations between the different sets of explanatory variables entering the export demand equations, the table below shows that the correlations between the relative prices are lower than those corresponding to the import demand equations (RP_i stands, as before, for the relative prices of the ith SITC section). As far as the activity

TABLE 3

CORRELATION COEFFICIENTS BETWEEN RELATIVE EXPORT PRICES

	RP_0	RP_1	RP_2	RP_5	RP_6
RP_0	1.000	0.262	0.315	0.030	0.385
RP_1		1.000	-0.053	-0.234	0.364
RP_2			1.000	0.275	-0.002
RP_5				1.000	-0.376
RP_6					1.000

variable is concerned, the index of industrial production of the O.E.C.D. countries has been introduced in all our export demand equations. The index of chemical production of the O.E.C.D. countries, which has been employed as activity variable only in the case of exports of chemicals, is highly correlated with the index of total industrial production (r=0.997).

Finally, lagged values of the dependent variables appear among the regressors, in two out of five, export demand equations (i.e. export demand equation for food and export demand equation for manufactures). Thus, in contrast to the case of imports, the argument concerning correlation between the different sets of explanatory variables in the case of exports is not valid.

In view of the above and taking into consideration the observed contemporaneous and lagged correlations between the disturbances, there is some evidence that the joint estimation of the selected forms of the export demand equations, might yield more efficient estimates. Further, it follows from the latter that the omitted common factors that affect the behaviour of the above groups of exported commodities, are also expected to cause serial correlation of the disturbances; such serial correlations have already been taken into account in the specification and estimation of the export demand equations. It appears then, that in the system of equations (1), apart from $\underset{\sim}{\Sigma}$, we can also treat the matrices $\underset{\sim}{R}_i$ as non-diagonal. That is, to assume that the cross-equation serial correlations of the disturbances can be represented by a general vector autoregression. In this case the set of equations in (1) can be estimated by the autoregressive maximum likelihood method (see Hendry (1971)). Moreover, a vector generalization of the model selection procedure based on the likelihood ratio test principle, already described in chapter IV, section 1.1., can be applied in order to specify the dynamic structure of the equations. However, such an approach was unattainable, since the relevant computer program (see Hendry and Srba (1978b)) required enormous process time for the estimation of models such as (1) with $n \geq 2$ and $\ell > 4$.

In conclusion, we take into consideration that on the one hand we are confronted by the above difficulties in estimation when $\underset{\sim}{R}_i$ are treated as non-diagonal, and on the other, there is the fact that the observed contemporaneous and lagged correlations between the residuals are quite small (though significant). Therefore, we believe that any loss in efficiency yielded from the estimation of the export demand equations by the (linear or non-linear) least-squares method is negligible, as well as that the already estimated preferred forms should be retained.

CHAPTER VIII

APPLICATIONS OF THE EMPIRICAL FINDINGS

The preceding chapters dealt with the estimation of the structural parameters of Greece's import and export demand equations. In the present chapter we attempt some applications of these estimates to the trade balance of Greece. Specifically, we try to predict the country's balance of trade for the year 1982, under different assumptions about growth rates in Greece and its trading partners, and relative price changes; thus, we show the sensitivity of the country's trade balance to these various assumptions.

1. Ex-post Predictions for 1977 and 1978

Before presenting any ex-ante predictions, it is interesting to test the performance of the estimated equations and to see how well they can explain economic behavior outside the confines of the sample. It should be mentioned, however, that a good performance of the estimated equations outside the sample period does not guarantee good ex-ante predictions for the forecast period, because the values of the explanatory variables in the prediction period are usually not known at the time forecasts are made.

The post-sample parameter stability test was applied to all import and export demand equations. The observations used for this purpose are the monthly figures for the year

1977, which meanwhile became available. If f_i denotes the forecast error and s^2 is the residual variance of the particular import or export demand equation, then the parameter stability test-statistic is $\frac{1}{s^2} \sum_{i=1}^{12} f_i^2$ which, under the null hypothesis that all parameters are stable and consistently estimated, is distributed asymptotically as a $X^2(12)$ variate. The table below shows the values of the stability test statistics of the preferred forms. Note that with addition of the monthly observations for 1977, the seasonal adjustment of the data was made for the entire sample before the above test was applied.

TABLE 1

POST SAMPLE PARAMETER STABILITY TESTS

(1977 Monthly Data)

Import Demand Equation for:	$X^2(12)$	Export Demand Equation for:	$X^2(12)$
(i) Food	11.848	(i) Food	7.159
(ii) Raw Materials	7.620	(ii) Beverages and Tobacco	32.299*
(iii) Fuels	18.329	(iii) Raw Materials	26.280*
(iv) Chemicals	8.076	(iv) Chemicals	14.135
(v) Manufactures	18.720	(v) Manufactures	4.801
(vi) Machinery and Transport Equipment	6.257	(vi) Overall Export Demand Equation	7.440
(vii) Overall Import Demand Equation	7.050		

*Significant at the 1% level

(The 5% and 1% critical values of the $X^2(12)$ distribution are 21.03 and 26.22 respectively)

It can be seen from the above table that in all equations, except the export demand equations for beverages and tobacco, and raw materials, the post-sample parameter stability test took a non-significant value. As fas as exports of beverages and tobacco are concerned, the above significant value of the post-sample parameter stability test is not a surprise if we take into consideration the erratic growth path of this export group over the sample period.

Before we present the ex-post predictions, we should mention that the use of monthly data in the ex-ante predictions, presented in the next section, caused some methodological difficulties. In particular, because of the seasonal variation of the data it is difficult to generate a sequence of monthly values of the explanatory variables for the future. Even if this was possible, our equations do not explain the seasonal variation in the dependent variable since, for the reasons described in chapter IV section 1.3., it was impossible to include seasonal dummy variables in the model. In the context of estimation from seasonally adjusted data, a possible solution could be to ignore the seasonality and generate monthly values of the explanatory variables, using different assumed trends. However, experiments made, applying this procedure, produced unacceptable results, since this approach requires a fairly big number of dynamic forecasts (see next section).

In view of the above we decided to use annual average data for our ex-ante predictions. That is, we generated a series of annual average values of the explanatory variables assuming different rates of annual growth. Then, applying to these series of data the preferred forms of our import and export demand equations we obtained the ex-ante predictions. First, in order to assess the performance of this procedure, we calculated ex-post predictions of Greece's imports and exports of goods for the first two post-sample years 1977 and 1978, using average annual data for these two years (for 1978 only the annual figures are available). The equations used for the ex-post predictions, are the preferred forms already described in chapters V and VI. Since two different model selection procedures have been applied for each import and export category, whenever there is more than one preferred form, both equations were used and the one which performed better was chosen (the post-sample parameter stability tests previously reported, were applied to these forms). As mentioned above, the actual values of the dependent and explanatory variables used in the ex-post predictions are average annual data for the years 1977 and 1978. That is, they are the averages of the corresponding monthly data for these two years and therefore they do not differ much from the values the variables take in the sample period.

It can be seen from the table below that all equations, except the export demand equation for beverages and tobacco, perform reasonably well in the post-sample period. For 1977, these equations have predicted Greece's trade balance with an error of -0.63 percent (excluding these groups of imported and exported commodities for which, due to data limitations, import or export demand functions have not been estimated). For 1978, Greece's trade balance has been predicted by an error of,

TABLE 2 -

ACTUAL AND PREDICTED VALUES OF GREECE'S IMPORTS AND EXPORTS OF GOODS FOR 1977 AND 1978
(In million Drachmas at 1970 prices)

Major Groups	1977			1978		
	Actual Value	Predicted Value	Errors as a Percent of Actual	Actual Value	Predicted Value	Errors as a Percent of Actual
IMPORTS OF:						
(i) Food	6960	7304	-4.94	8117	7500	7.60
(ii) Raw Materials	7071	7075	-0.06	7160	7017	2.00
(iii) Fuels	5915	6466	-9.32	7737	6955	10.11
(iv) Chemicals	8179	8550	-4.54	8269	8905	-7.69
(v) Manufactures	15482	14731	4.85	16043	16181	-0.86
(vi) Machinery and Transport Equipment[1]	25718	24021	6.59	25134	23440	6.74
Total (i-vi)	69325	68147	1.70	72460	69998	3.40
Total (from the over-all equation)	73474	72323	1.57	77201	76412	1.02
EXPORTS OF:						
(i) Food	8794	8539	2.90	9382	8317	11.35
(ii) Beverages and Tobacco	3505	3024	13.72	4883	3237	33.71
(iii) Raw Materials	3170	3092	2.46	3424	3149	8.03
(iv) Chemicals	3054	2658	12.97	3019	2924	3.15
(v) Manufactures	12999	13269	-2.08	13021	13601	-4.45
Total (i-v)	31522	30582	2.98	33729	31228	7.41
Total (from the over-all equation)	42831	40225	6.08	49443	43319	12.39
TRADE BALANCE						
(i) From Disaggregated Equations	-37803	-37565	-0.63	-38731	-38770	0.10
(ii) From the over-all Equations	-30643	-32098	4.75	-27758	-33093	19.22

[1] Excluding ships

only, 0.10 percent.

Prediction from the component equations for merchandise imports is slightly worse than that from the overall equation. This, of course, does not mean that there is little to be gained by working with the disaggregated equations even from the point of view of prediction. The advantage of disaggregation is that commercial policies and other structural changes can be more readily incorporated into the disaggregated system. Moreover, with the disaggregated form we are able to obtain adittional information, that is, to introduce and evaluate the importance of factors which are submerged at the aggregate level.

On the other hand, the prediction from the component equations for merchandise exports is clearly better than that from the overall export demand equation. However, these predictions are not quite comparable since the groups of exported commodities "excluded" from the analysis account for about 29 percent of total exports during the period 1977-1978 (in the case of imports the omitted categories account for only 6 percent of total imports). As mentioned in chapter VI section 1., these omitted commodity groups, being unimportant in Greek export trade, are not listed for the period before 1972. However, since 1972 they started to have a noticeable share in total exports. About 50% of these commodity groups consists of exports of machinery and transport equipment as well as miscellaneous manufactures. The latter groups of exports appear to have a non-erratic path of growth. However, the remaining 50% of these "omitted" commodity groups, consists of exports of fuels and lubricants, and these exhibit a highly erratic growth path over time. Therefore, in view of data limitation, it is difficult to forecast the growth of the above three groups taken together.

Finally, it appears from the above table that all equations, except the import demand equation for manufactures, the overall import demand equation and the export demand equation for chemicals, perform better in 1977 than in 1978.

It should be mentioned that the preferred forms of the import and export demand equations, which were used for the predictions, are in log-linear form, and so, the forecasts were obtained from the antilogs of the predicted values. This introduces an upwards bias since, though in the log-linear form the predictor is unbiased, the antilog of the predictor is not. If the predictor $\hat{\log}(Y)$ is normally distributed with mean $E\{\hat{\log}(Y)\}$ and variance σ^2, then $\exp\{\hat{\log}(Y)\}$ is log-normally distributed with mean

$$\begin{aligned} E\{\exp[\hat{\log}(Y)]\} &= \exp\{E[\hat{\log}(Y)] + \tfrac{1}{2}\sigma^2\} = \\ &= \exp\{\log(Y)\}\cdot\exp(\tfrac{1}{2}\sigma^2) = \\ &= Y\, e^{(1/2)\sigma^2} \end{aligned}$$

where $\log(Y)$ denotes the value of the prediction from the log-linear form with known population parameters. Thus, taking the series expansion for $e^{(1/2)\sigma^2}$, the bias is

$$E\{\exp[\hat{\log}(Y)]\} - Y = Y(\tfrac{1}{2}\sigma^2 + \tfrac{1}{4}\tfrac{\sigma^4}{2!} + \tfrac{1}{8}\tfrac{\sigma^6}{3!} + \cdots)$$

Hence, the transformation from logarithms to levels has increased the forecast errors as a percentage of the actual values. In particular, the percentage errors in logarithms were about four times smaller than those in levels which are presented in the above table.

The values of the post-sample parameter stability test statistics, as estimated from the above ex-post predictions, are shown in the table below. Since the data used are annual averages, the asymptotic variance of the forecast error is 1/12 of that in the case of monthly data. Thus, the $X^2(2)$ test-statistics computed from the relevant programs have been multiplied by 12.

TABLE 3

POST SAMPLE PARAMETER STABILITY TESTS

(1977 and 1978 Annual Average Data)

Import Demand Equation for:		$X^2(2)$	Export Demand Equation for:		$X^2(2)$
(i)	Food	2.342	(i)	Food	5.136
(ii)	Raw Materials	0.198	(ii)	Beverages and Tobacco	9.434*
(iii)	Fuels	1.054	(iii)	Raw Materials	2.875
(iv)	Chemicals	2.861	(iv)	Chemicals	1.262
(v)	Manufactures	1.970	(v)	Manufactures	0.408
(vi)	Machinery and Transport Equipment	5.388	(vi)	Overall Export Demand Equation	9.576*
(vii)	Overall Import Demand Equation	4.522			

*Significant at the 1% level

(The 5% and 1% critical values of the $X^2(2)$ distribution are 5.99 and 9.21 respectively)

We can see from the above table that in all equations, except the export demand equation for beverages and tobacco, and the overall export demand equation, the post-sample parameter stability test (for annual average data) took a non-significant value.

In summary, with the exception of the fairly large error in the predicted value for the exports of beverages and tobacco in 1978, all equations fit the (annual) data of the post-sample period quite well, and this suggests acceptance of the model for ex-ante predictions.

2. Ex-ante Predictions for 1982

Predictions from an econometric model are conditional forecasts and their validity depends on the realism and relevance of the assumptions on which they are based.

It is assumed that the structural equations observed in the past will continue to hold into the future, that is, no structural changes take place between the observed and future period. Generally, the postulate of an unchanged structure is highly plausible for relatively short-term predictions.

In the absence of specification error in the estimated equation, we may distinguish three possible sources of forecast error. In the first place, the values of the explanatory variables in the prediction period are usually not known at the time forecasts are made. The explanatory variables, namely the activity and relative price variables of this study, are assumed to have certain values. If these assumptions are not valid, our predictions cannot be regarded as estimates of Greece's trade balance. For these reasons, alternative values for the explanatory variables were used and various predictions under these hypotheses were made. Therefore, range estimates rather than single estimates of Greece's trade balance are calculated, revealing the sensitivity of the country's balance of trade to various alternative values of the activity and relative price variables. As was mentioned above, such forecasts are conditional forecasts, showing alternative estimates of alternative assumptions.

The second source of forecast error is the contribution of the error in the estimation of the parameters of the forecast equation and reflects the random variation of the sample period disturbances. The parameters of the forecast equation are only estimates of the population parameters and, since they are based on finite samples, they are subject to sampling errors.

The third source of forecast error arises from the disturbance term in the forecast period. This is set equal to zero for forecasting purposes but its actual value will be different from zero. Thus, improvements in estimation will reduce the (conditional) forecasting error but this error cannot be reduced below the disturbance term in the prediction period (see Wallis (1979), ch. 4, for a fuller treatment of prediction problems).

A further difficulty in forecasting arises from the dynamic form of our import and export demand functions. The preferred forms of our equations are either of the original equation, autoregressive error specification form, where dynamics are introduced through error dynamics, or of the unrestricted transformed equation form, where dynamics are expressed through lagged values of the dependent and explanatory variables. In few cases we have a mixture of dynamics, that is, both systematic and error dynamics. Thus, in calculating a sequence of forecasts, using the above equations, the forecast for one or more periods enter into the calculation of the next (if error dynamics are introduced into the equation, best linear unbiased predictions are obtained from the corresponding restricted transformed equation).

In order to obtain predictions of Greece's imports and exports for the year 1982, we have generated a sequence of annual (average) values of the explanatory variables entering our import and export demand equations, for the period 1977-1982 according to a given rate of annual change. Then we apply to these series of data the preselected preferred form of the particular import or export demand equation. However, in this way a model with a dynamic structure describing monthly movements, is applied to annual data.

To avoid this, two alternative procedures might be used. Firstly, instead of generating a sequence of annual values of the explanatory variables, we can generate a sequence of monthly values. For instance, if r is the assumed annual rate of growth of the activity variable, we can generate a sequence of monthly data for this variable using the monthly rate of growth $s = (1+r)^{1/12}-1$. Then, instead of predicting an annual index of volume of imports (or exports) for 1982 we predict 12 monthly indices and their average will be our prediction for that year. However, experiments made, applying this procedure, produced unacceptable results, since this approach requires a fairly big number of dynamic forecasts (i.e. 72) (for instance, as the index of industrial production was increasing the volume of imports (or exports) was either increasing at a rate lower than what was expected, or it was remaining almost unchanged; in some cases where the assumed annual rate of growth of the activity variable was small, i.e. 1% or 3%, the volume of imports (or exports) was decreasing rather than increasing).

An alternative approach is to obtain forecasts applying to the generated annual data, only the structural part of an import or export demand equation. That is, to use only the structural coefficients, ignoring the autocorrelation parameters. However, such an approach could be used only if for all import and export demand equations, the preferred form was the original (static) equation, autoregressive error specification.

A final point to be made is about the bias involved due to the transformation from logarithms to levels, as already mentioned in the previous section. Since the expected value of the antilog of the prediction equals to the true value of the prediction times $\exp\{(1/2)\sigma^2\}$, where σ^2 is the variance of the predictor in logs, the forecast of the trade balance is, in absolute terms, biased upwards. However, to the extend that the forecasts of the balance of trade presented below, aim only to indicate the relative position of the country's trade balance under different assumptions, the above defficiencies of our approach can be overlooked.

The table below shows Greece's balance of trade in 1982 (at 1970 prices) for various combinations of industrial production growth rates in Greece and O.E.C.D. countries. It is assumed that the relative prices prevailing in December 1976 (the last month of the sample period) will also prevail in 1982.

We should mention here, that for imports and exports of chemicals, predictions were obtained from the preferred forms of these equations, but where the indices of total industrial production of Greece and the O.E.C.D. countries respectively, were used as activity variables instead of the corresponding indices of chemical production which were initially employed. We also recall, that the groups of exported commodities excluded from the analysis accounted for about 25 percent of total exports during the last three years of our sample period. Thus, on the assumption that they will continue to contribute the same percentage to total exports in the future, to allow for these "excluded" exports, the sum of the predicted values of exports from the disaggregated functions has been increased by 33.3 percent.

The main conclusions that can be drawn from the above predictions may be summarized as follows.

(i) Each 2 percent increase (decrease) in the annual rate of growth in Greece's industrial production deteriorates (improves) the country's trade balance by 7,558 million dracmas on average (this marginal deterioration of the trade balance increases along with the Greek growth rate)

(ii) For each 2 percent decline (rise) in the annual rate of growth in the industrial production of the OECD

TABLE 4

GREECE'S TRADE BALANCE IN 1982, ASSUMING DIFFERENT COMBINATIONS OF AVERAGE ANNUAL GROWTH RATES BETWEEN 1976 AND 1982 FOR INDUSTRIAL PRODUCTION IN GREECE AND O.E.C.D. COUNTRIES

(In million Drachmas at 1970 prices)

		O.E.C.D. Growth Rate				
	%	1	3	5	7	9
	3	-28928	-22636	-15444	- 7233	+ 2138
Greek Growth Rate	5	-35903	-29611	-22419	-14208	- 4837
	7	-43446	-37154	-29962	-21751	-12380
	9	-51602	-45310	-38118	-29907	-20536

countries, Greece's balance of trade deteriorates (improves) on average by 7,767 million drachmas (this increment in the deterioration of trade balance increases along with the OECD growth rate).

(iii) If the annual rates of growth of both the Greek and O.E.C.D. industrial productions increase (decrease) by 2 percent, the country's trade balance improves (deteriorates) by an average of only 209 million drachmas. On the other hand, uniform rates of growth in both the Greek and O.E.C.D. industrial productions improve Greece's balance of trade by an average of only 700 million drachmas. The above result from the fact that the income elasticity of domestic demand for commodity imports divided by the export-import ratio is slightly lower than the income elasticity of foreign demand for Greek goods (for the period 1954-1976, excluding imports of ships, the export-import ratio amounted to 0.462, whereas, using the average shares, the total income elasticities of domestic and foreign demand for imports and exports, are 0.847 and 1.977 respectively). Thus, at constant prices, the outcome of the country's balance of trade depends on the income growth rates of Greece and O.E.C.D. countries, the income elasticities of imports and exports, as well as on its export-import ratio (see Prodromidis (1974), p p. 32-36).

Let us now try to assess the sensitivity of Greece's balance of trade to changes in relative prices. The table below shows the predicted values of Greece's trade balance in 1982 for various combinations of import and export relative price changes, but at growth rates of 7 and 3.3 percent in the industrial production of Greece and O.E.C.D. countries respectively. These growth rates are slightly lower than those experienced during the 1970-1976 period. As far as Greece's industrial production is concerned, the assumption of an annual rate of growth of 7 percent is consistent with the five year (1978-1982) economic development plan for Greece (Center of Planning and Economic Research, Social and Economic Development Plan for Greece 1978 - 1982, Athens, 1978). On the other hand, the annual rate of growth of 3.3 percent in the industrial production of the O.E.C.D. countries, is the average of the realised one in the period 1976-1978 and the one suggested for the period 1978-1982 by the Centre for Economic Forecasting of The London Business School (Economic Outlook 1979-1983, Forecast Release, Volume 4, Number 1, October 1979, p. 1).

The major findings that emerge from these predictions may be summarized as follows:

(i) Each 2 percent increase (decrease) in the annual rate of change in relative import prices improves (deteriorates) the country's trade balance by 9,084 million drachmas on average (this marginal improvement of the trade balance declines as the annual rate of change in relative import prices increases)

(ii) Each 2 percent increase (decrease) in the annual

TABLE 5

TRADE BALANCE OF GREECE IN 1982 UNDER DIFFERENT RATES OF CHANGE IN RELATIVE IMPORT AND EXPORT PRICES AND FOR 7% AND 3.3% RATES OF GROWTH OF INDUSTRIAL PRODUCTION IN GREECE AND O.E.C.D. COUNTRIES RESPECTIVELY

(In million Drachmas at 1970 prices)

		Annual Rate of Change in Relative Import Prices									
	%	-9	-7	-5	-3	-1	1	3	5	7	9
Annual Rate of Change in Relative Export Prices	-9	- 76435	-62150	-49641	-38658	-28977	-20416	-12825	- 6077	- 61	+ 5318
	-7	- 81234	-66949	-54440	-43457	-33776	-25215	-17624	-10876	- 4860	+ 519
	-5	- 85503	-71218	-58709	-47726	-38045	-29484	-21893	-15145	- 9129	- 3750
	-3	- 89322	-75037	-62528	-51545	-41864	-33303	-25712	-18964	-12948	- 7569
	-1	- 92750	-78465	-65956	-54973	-45292	-36731	-29140	-22392	-16376	-10997
	1	- 95839	-81554	-69045	-58062	-48381	-39820	-32229	-25481	-19465	-14086
	3	- 98633	-84348	-71839	-60856	-51175	-42614	-35023	-28275	-22259	-16880
	5	-101170	-86885	-74376	-63393	-53712	-45151	-37560	-30812	-24796	-19417
	7	-103483	-89198	-76689	-65706	-56025	-47464	-39873	-33125	-27109	-21730
	9	-105597	-91312	-78803	-67820	-58139	-49578	-41987	-35239	-29223	-23844

rate of change in relative export prices deteriorates (improves) Greece's balance of trade by an average of 3,240 million drachmas (this increment in the deterioration of trade balance declines as the annual rate of change in relative export prices increases).

(iii) From the above it results that a 2 percent increase (decrease) in the annual rate of change in relative import prices and a 2 percent decrease (increase) in the annual rate of change in relative export prices, improve (worsen) the country's balance of trade by an average of 12,324 million drachmas.

(iv) If the annual rates of change in relative import and export prices increase (decrease) by 2 percent, the country's trade balance improves (deteriorates) by an average of 5,843 million drachmas. This suggests two things; first, the relative price elasticity of import demand for goods is sufficiently higher than that of export demand for goods (see chapter V, section 2.7. and chapter VI, section 2.6.); and second, these price elasticities are large enough for a devaluation to improve the country's balance of trade.

In summary, our analysis indicates that, other things being equal, disparities in the rates of inflation in Greece and its trading partners cause (through the changes in relative prices) substantial changes in the country's trade balance. It appears, also, that the increase (decrease) in the annual rate of change in relative export prices, should be, on average, 2.8 times the increase (decrease) in the annual rate of change in relative import prices, if Greece's trade balance is not to deteriorate.

During the period 1970 - 1978, export prices have increased at an annual rate approximately equal to that of Greece's domestic prices, whereas the annual rate of growth of import prices is about 1.4 times the average annual rate of growth of the domestic prices of the O.E.C.D. countries (similar relations hold for the whole sample period). If these relations between the increases of the prices will hold approximately in the future, Greek prices should rise at a rate not more than 20 percent of that in the O.E.C.D. countries, if the deterioration (in constant prices) of Greece's balance of trade is to be avoided.

CHAPTER IX

CONCLUSIONS

In the preceding chapters we described model selection procedures for the empirical specification of dynamic models, and an attempt was made, applying these procedures, to obtain the preferred specifications and their numerical estimates of the import and export demand functions of Greece. We, also, tried to assess the effects on Greece's trade balance at various growth rates in Greece and the O.E.C.D. countries, and different relative price changes.

In section 1. of this final chapter we discuss the experience gained from the application of the above model selection procedures in the context of a large scale study as the present one. In section 2. we deal briefly with the implications of the empirical findings for Greek trade policy.

1. Conclusions from the Application of the Model Selection Procedures

Thus far, we have tried to obtain the preferred specifications and their numerical estimates for thirteen import and export demand equations of Greece. As already mentioned in chapter IV, two model selection procedures have been applied for each import and export demand equation. In the first procedure we start with the simplest (static) model and we test if it is necessary to consider a more general one, whereas in the second we begin with a general unrestricted dynamic model and then an attempt is made to reduce the number of parameters needed to specify the data generation process.

In many cases the application of the two different model selection procedures has resulted in different dynamic specifications. To facilitate the presentation of the various preferred specifications, we tabulate the orders of dynamics of all the import and export demand equations we have reported earlier in the chapters V and VI. In particular, the table below shows the order of the systematic and error dynamics as suggested by the likelihood ratio, Wald and difference Wald criteria, for each import and export demand equation.

The "common factor" analysis has been applied to both the a priori specified unrestricted dynamic model and the simplified one (by testing for zero roots from the set of r common roots initially extracted from the general unrestricted dynamic model). Thus, with regard to the Wald and difference Wald criteria, the orders of dynamics suggested by the application of these criteria in both the general and simplified unrestricted dynamic models are shown in the table below. We also recall that when the Wald or the difference Wald criteria suggested that a structural equation form with lagged values of the variables should be considered, then the error dynamics were determined after this structural equation was estimated subject to an autoregressive error term, and in likelihood ratio tests the preferred autoregressive error form was selected.

A number of interesting points arise from this table. In only two cases did all the methods lead to the same specifi-

TABLE 1

DYNAMICS OF THE PREFERRED SPECIFICATIONS FROM THE APPLICATION OF THE MODEL SELECTION PROCEDURES

Equation	Likelihood Ratio Tests			Wald Criteria						Difference Wald Criteria					
				General UTE			Simplified UTE			General UTE			Simplified UTE		
	q	m	q-m	q	m	q-m	q	m	q-m	q	m	q-m	q	m	q-m
M 0	9	9	0	9	9	0	9	9	0	6	5	1	6	5	1
M 2	11	11	0	11	11	0	11	11	0	13	12	1	13	12	1
M 3	12	12	0	12	12	0	12	12	0	13	12	1	13	12	1
M 5	12	12	0	12	12	0	12	12	0	13	12*	1	13	12*	1
M 6	4	4	0	4	4	0	4	4	0	4	3*	1	4	4	0
M 7	3	0	3	4	3	1	4	3	1	5	3	2	4	3	1
M 0-9	9	9	0	9	9	0	9	9	0	10	9	1	10	9	1
X 0	13	1	12	13	13*	0				13	12*	1			
X 1	4	4	0	4	4	0	4	4	0	4	4	0	4	4	0
X 2	2	2	0	2	2	0	2	2	0	2	0	2	2	2	0
X 5	6	6	0	6	6	0	6	6	0	6	6	0	6	6	0
X 6	7	0	7	7	7*	0	7	7*	0	-	-	1	-	-	1
X 0-9	12	0	12	12	12*	0	12	12*	0	-	-	1	-	-	1

Definitions: q is the overall dynamics of the equation
m is the error dynamics
q-m is the systematic dynamics
M i (X i) is the import (export) demand equation for the ith SITC group of traded commodities
M 0-9 (X0-9) is the aggregate import (export) demand equation

* the autoregressive error hypothesis is rejected in favour of the unrestricted transformed equation

\- an equation was not estimated because of convergence difficulties cation (equations X 1 and X 5). Except for the difference Wald criteria, as used in the general unrestricted model, the other criteria suggested the same order of systematic and error dynamics in only 2 cases (equations M 6 and X 2). Notice the relatively low order of the overall dynamics of these preferred specifications.

In 9 out of 12 cases (equation X 0 is excluded) the application of the difference Wald criteria suggested the same order of systematic dynamics in both the general and the simplified unrestricted dynamic models. Notice that in those 3 cases where the difference Wald criteria suggested different systematic dynamics (i.e. equations M 6, M 7 and X 2), the maximum number of lags in the simplified unrestricted equation was much lower than that of the general dynamic model. Also, in all the above 12 cases, the Wald criteria led to the same specification regardless of whether they were applied to the general form or the simplified form of the unrestricted dynamic equation.

The Wald and difference Wald criteria, applied to the general unrestricted dynamic model, suggested the same factorization of the overall dynamics into error and systematic dynamics, in only two cases (equations X 1 and X 5). But, when they were applied to the simplified version of the general dynamic model, they suggested the same specification in 5 cases (equations M 6, M 7, X 1, X 2 and X 5). Notice that in these 5 cases the corresponding simplified unrestricted models have the lowest overall dynamics comparing with the other equations. Finally, the likelihood ratio and the Wald criteria suggested the same order of systematic and error dynamics, in 9 out of 13 cases.

Considering now the kind of dynamics suggested from the above criteria, we can see from the above table, that in all cases except one (equation M 7) the Wald criterion had a non-significant value suggesting that the overall dynamics in the equation should be expressed only through error dynamics. With regard to the likelihood ratio tests, the autoregressive error hypothesis was accepted against the unrestricted transformed equation in 9 out of 13 cases. It appears then, that there is a tendency in the common factor analysis to reject the common factor restrictions too infrequently (see also Hendry and Mizon (1978)). This is related to another revealing result of our empirical analysis, that is the Wald criterion took in all cases a value smaller than that of the likelihood ratio. This contradicts what has been established for general linear constraints in the multivariate linear regression model - that $W \geqq LR$. In the context of the dynamic specification problem with first order error dynamics, Mehta (1979) examined, in a simulation study, the behaviour of these two tests and found the same relationship between the Wald and likelihood ratio criteria for non-linear restrictions (i.e. $W < LR$). (see also Savin (1976) and Berndt and Savin (1977)).

Another point to be mentioned is the overall dynamics of the equation in those cases where the difference Wald criteria suggested that a structural equation form with lagged values of the variables should be considered. In most of these cases, when the new structural equation was reestimated subject to an autoregressive error term, the selected autoregressive error specification had the same (or higher) order as the initially selected autoregressive error form, where no systematic dynamics were introduced into the equation (e.g. equations M 2, M 3, M 5, M 7, M 0-9). That probably reflects the fact that, in general, there is a one-to-one correspondence between ℓ periods lagged values of all the variables and ℓth order error dynamics, so that even though lagged values of all the variables (of order lower than ℓ) are introduced into the equation, the order of error dynamics remains the same. In other words, it seems that the order of the overall dynamics is increased rather than remaining the same as it should be expected from its factorization into systematic and error dynamics.

It also appears that the Wald difference criterion is more sensitive than the Wald criterion to the maximum number of lags included into the general unrestricted model. That is, as the order of overall dynamics in the general unrestricted dynamic model is reduced, due to its simplification, it is more likely that the difference Wald criterion will take smaller values than the ones previously obtained (e.g. equations M 6, M 7 and X 2) - i.e. it suggests systematic dynamics of lower order than that initially suggested when it was applied to the general unrestricted dynamic model.

Finally, in the context of estimation from monthly data, our empirical analysis gives rise to the following points. The use of monthly data implies that (systematic or error) dynamics of high order (i.e. 13 or even more) should be considered and this may be the cause of the following problems.

First, it was observed that the higher the order of dynamics, the bigger the difference between the computed Wald and likelihood ratio criteria (with likelihood ratio criterion being persistently bigger). However, this increasing difference between the two criteria might be due to the increasing number of degrees of freedom and therefore to the increasing variance of the asumptotic X^2 distribution. To study this we computed the average difference between the two criteria for all the orders of dynamics (which, as mentioned in chapter IV, we have considered for every import and export demand equation, but not reported), i.e. for m = 1, ... , 13, where m denotes the order of error dynamics. Then, we divided these average differences by $\sqrt{2m}$ (the standard deviation of the asymptotic X^2 distribution) and we found that this ratio was increasing slightly along with the order of dynamics. This indicates that there is a tendency for the (actual) difference between these two tests to increase slightly more than what one would expect from using the number of the degrees of freedom (and therefore the variance of the asymptotic X^2 distribution) as a rough guide .

Mehta (1979) found that, under the null hypothesis that a common factor polynomial of degree one exists, compared to the asymptotic X^2 distribution, the Wald criterion is downward biased whereas the likelihood ratio criterion is upward biased. That is, the size of the Wald test was smaller than the nominal one whereas the size of the likelihood ratio test was bigger.

Our results indicate that, under the null hypothesis that a common factor polynomial of degree m exists, the diversion of the Wald or likelihood ratio criterion (or both) from the asymptotic (central) X^2 distribution increases along with the number of degrees of freedom (i.e. the order of dynamics).

Of course this indication is based on the assumption that in the cases which we have examined the null hypothesis is true; but this is not known (though in most cases both the Wald and the likelihood ratio test had a non-significant value). If the null hypothesis is not valid, then the observed increasing difference between the two tests, can be interpreted as an increasing difference between the powers of the two tests along with the increasing order of dynamics. If Mehta's (1979) results, mentioned above, are also valid for higher order dynamics, then in view of our results it seems that as the order of dynamics increases, compared to the nominal size of the test, either the size of the Wald test decreases or the size of the likelihood ratio test increases (or both).

Of course, it is difficult to say which of these two tests has a size closer to the nominal size of the test (as well as which is the exact direction of the diversion from the nominal size of the test) in cases of higher order dynamics (for first order dynamics, Mehta (1979) found that under the null hypothesis the distribution of the Wald criterion is closer than that of the likelihood ratio criterion to the asymptotic X^2 distribution). A simulation study is required in order to illuminate this point and also to examine, under the alternative hypothesis, the power of these tests. Hence, in view of the above it appears that for higher order dynamics (and therefore for a bigger number of degrees of freedom) the actual size of these tests may differ more than in lower order dynamics from the nominal size of the test.

Secondly, another problem is that estimating autoregressive error forms of high order the problem of multiple minima becomes more serious: whenever we failed to obtain acceptable results the general autoregressive error specification was of order higher than three. Thirdly and finally, in addition to the computational expense of estimating autoregressive error forms of high order, we are restricted not to include seasonal dummy variables in the model (see chapter IV, section 1.3.) if we are to estimate the seasonal variation within the model.

In conclusion we can say that the two procedures should be considered as supplementary rather than competitive, since whatever is the indication from the common factor analysis, a non-linear estimation of the suggested specification will follow and a likelihood ratio test for the discrimination between the restricted and unrestricted forms will too (in fact this test is automatically computed from the relevant computer programs). Of course, as far as the power properties of the sequential testing procedure (i.e. common factor analysis) are concerned, this test of the accepted hypothesis against the maintained hypothesis is irrelevant. The optimal procedure for an ordered nested structure is the sequential one without the, so called, final check (see Mizon (1977), p. 1226).

Thus, the question arising is whether in practice the researcher will be restricted either to the Wald testing principle or to the likelihood ratio testing principle, or can he use both in an effort to select the model most consistent with the data. In view of our observed discrepancies between the selected models from the two procedures it becomes apparent that one will be tempted to check a decision from one procedure against that from the other seeking, as it were, a second opinion. In view of the former discussion, such an approach does not seem to be unreasonable in the sense that if both tests take a significant (or non-significant) value then one feels more secure that the decision reached is correct. If, however, the two tests contradict each other (i.e. the Wald test accepts the autoregressive error hypothesis whereas the likelihood ratio test rejects it) then the choice between the two models could be done on other grounds (e.g. post-sample parameter stability test). Moreover, the adequacy of a model suggested by, say, the common factor analysis, is tested using various diagnostic tests (e.g. Box-Pierce test statistic for random correlogram, etc.). However, these tests are rather tests of misspecification and are related to the "specific-to-general" modelling approach.

In view of this discussion it becomes apparent that the performance of these criteria should be exploited further. In particular, further research might be worthwhile along the following lines. First, more applications of the Wald and difference Wald criteria to the dynamic specification problem are necessary, in order to gain more experience on their performance in comparison with that of the likelihood ratio testing principle. Second, further Monte Carlo studies on the behaviour of the above criteria are required. Specifically, it will be worthwhile to look at cases where the true order of systematic and error dynamics of the relationship under study is larger than 1, paying attention to the difference Wald criteria. For the latter criterion it will be interesting to examine its performance in relation with different number of maximum lags in the unrestricted dynamic model.

2. Implications for Trade Policies

Our predictions in the previous chapter indicated a tendency towards a deterioration in the country's balance of trade in the years to come. In particular, under the basic assumptions of 7 and 3.3 percent growth rates in the industrial production of Greece and O.E.C.D. countries respectively, the trade balance deficit is expected to exceed the amount of 35,000 million drachmas (at 1970 prices) in 1982 (see chapter VIII, table 5).

However, it appears that, in terms of trade imbalances, the performance of the Greek economy could be more encouraging in the further future. This is partly supported by our empirical results. In particular, it was found in this study, that the income elasticity of foreign demand for Greek goods is substantially greater than the income elasticity of domestic demand for commodity imports. On the contrary, we found that imports are, in general, more sensitive to changes of relative prices than exports. These two taken together, are expected to have a favorable effect on the trade balance of Greece in the future.

On the other hand, the efforts of Greece to change the commodity composition of exports in favor of goods of high income elasticity, such as chemicals, manufactures, etc., has met with success. This was the result of a number of measures taken in the sixties by the government, which intended to industrialize the country and promote commodity exports. The consequence of this policy was a shake-up of the traditional economic structure of the country (until the early sixties Greek economy was based primarily on the agricultural sector) manifested by the construction and operation of several large plants throughout the country. Thus, exports of chemicals and basic metal products, which were unimportant in the earlier years of the sample, became quite impressive in later years. Furthermore, the prospects for an accelerated industrialization of the country in the next ten to fifteen years appear to be brighter than past achievements seem to indicate.

There are, however, some obstacles which may hinder improvements in the country's trade balance in future. The income elasticities of the import demand for raw materials, fuels and machinery and transport equipment were found to have a value higher than one. At the same time, the import demand for raw materials and fuels was found to be price inelastic. These indicate that industrialization in Greece generates a direct import demand for raw materials, fuels and capital equipment, and an attempt to interrupt the flow of these imports would undoubtedly destroy the country's efforts for economic development. Moreover, the recent price increases of raw materials and particularly of crude oil are expected to affect unfavorably the country's trade balance in the future. Thus, the process of economic development in Greece cannot be sustained without an adequate supply of foreign exchange to finance the rising import requirements of the country. It, therefore, appears that a relatively high rate of growth in Greece's export earnings is of strategic importance for its internal development.

Thus, Greece, a small country poor in soil and subsoil resources, is highly dependent on international trade and its strategy for economic development seems to be concentrated on the expansion of trade. In general, small countries, or countries with limited natural resources, are more likely to find that economic growth can be accelerated by international trade and, consequently, by international specialization. International trade, especially in small countries, intensifies competition, promotes technical efficiency and progress, and contributes in turn to rapid productivity advances and consequently to higher standards of living.

Greece's further export orientation is of vital importance, not only from the point of view of trade balance equilibrium, but also because it will set in motion a process of further technical progress and efficiency and it will provide the necessary conditions for a sustained process of development. It is apparent, therefore, that efforts should be made to increase and diversify the export capacity of the country.

In view of the above and the results of the preceding analysis, it appears that in order to improve the trade balance in the future, Greece may choose to pursue a policy aiming at the following:

(i) To try, in a positive way, to affect the income elasticity of foreign demand for Greek products. This could be attained by (a) improving the quality of exportable commodities through undertaking market research in foreign markets, and (b) advertising Greek products abroad.

(ii) Greater effort has to be made to change the composition of exports in favor of goods of high income elasticity, such as agricultural products, chemicals and manufactures. With regard to exports of agricultural products, there is a large market for fresh fruits and cotton in Europe. Therefore, it is proposed that the production of these more dynamic agricultural products be increased with a subsequent increase in their exports to follow. Such an increase in exports of the latter products will be at the expense of other traditionally exported agricultural commodities (tobacco, raisins, oil) for which the world demand is rather sluggish. On the other hand, the rising importance of chemical products, throughout the sample period, among the country's exportable commodities is an indication that Greece may become in the future a net exporter of chemicals in general. Finally, attention should be focused upon the further expansion of exports of manufactures. The export expansion of manufactures is important not only from the point of view of trade balance but also from the point of view of industrial development. Such a policy implies that an increase in capital investments in manufacturing is necessary for the export promotion of manufactures. Moreover, efforts should be made to reduce exports of raw materials per se (e.g. products of mining) or even semi-finished goods (e.g. basic metal products) in favor of more exports of finished manufactured goods which will help close, to a large extend, the trade gap of the country.

(iii) It was pointed in the previous chapter that different rates of inflation in Greece and its trading partners exert an important influence on the country's balance of trade. Thus, cost-wage and price policies should be designed to prevent Greek prices from rising faster than those of the country's main importers.

(iv) To control imports, given the income elasticities of demand for imports, Greece could choose to adopt a slower income growth. However, such a proposition is not desirable on other grounds.

(v) Finally, the above suggested policy measures would be substantially facilitated by goverment grants aiming at promoting research in general. That is, to develop better ways of exploiting the known natural resources of the country, to search throughout the country for new raw materials, etc.

It is obvious, that the above suggested policy measures aim to influence only those variables or parameters of the import and export demand equations, which can be affected by domestic policies. These variables or parameters which are exogenous to the Greek economic system (e.g. import prices, rate of growth of foreign income , price level of foreign substitutes and foreign income elasticities of Greek exportable commodities) and, consequently, cannot be influenced by domestic policies, depend on each time international economic conditions and Greece should adapt its trade policy to these conditions for a more favorable effect on its trade balance.

Greece must also continue its efforts to increase the rate of growth of invisible receipts. As mentioned in chapter II, international tourism and shipping have become important export industries and major sources of the country's foreign exchange earnings, financing about 25 percent of its merchandise imports. The expansion of these services will help to cover a great part of the trade deficit. Prospects are favorable both because world demand for tourist services is growing rapidly and because Greece is suited to supply these services (the income elasticity of foreign demand for tourism has been found to have a value higher than 2 - see Paraskevopoulos (1970), ch. VI). On the other hand the increasing share of Greek-controlled tonnage in the total world tonnage (6.1 percent in 1954, 12.0 percent in 1976) in relation to the expansion of world trade are both expected to contribute to the increase of the country's shipping exchange earnings.

A final point to be made concerns the effect on Greece's trade balance from the country's association with European Economic Community (the Greek accession treaty was signed in June 1979 for entry on January 1st 1981). Though the preceding analysis does not allow for clear-cut inferences about the trade relations of Greece with the E.E.C. countries (this can be the object of another study where separate import and export demand equations per country of origin in the case of imports, and of destination in the case of exports, will be specified, so that the special commercial relations of Greece with, say, the E.E.C. countries would be brought into focus), it is useful to mention briefly the prospects for Greece of joining the E.E.C..

In November 1962 agreement was reached with the European Economic Community whereby Greece became its first European associate member. Since July 1968 Greek industrial goods have been admitted to the original member states of the E.E.C. free of duty and certain agricultural products receive special treatment. Under an interim agreement signed in July 1975 the provisions of the 1962 aggreement were extended to the three new members of the E.E.C. (United Kingdom, Denmark and the Irish Republic). Since 1st July 1977 Greek industrial goods are granted duty-free entry into these three countries, and certain Greek agricultural goods are accorded duty-free treatment.

Under a financial Protocol between Greece and E.E.C. signed in 1977, Greece will receive a loan of 280 million European monetary units (i.e. about \$ 350 million) until 1981. The main productive sector which will be supported by these funds is agriculture with the purpose of modernizing it and increasing its productivity. It is planned to bring about changes in the structure of Greek agriculture with emphasis being laid on those crops which will not compete with European crops but will fill complementary needs of the Community.

In view of the above Greek agricultural products are expected to face a less risky export market and over all more optimistic economic conditions.

On the other hand, the technological assistance and funds for remodeling and updating the Greek productive machinery, provided by the community authorities, will contribute to the expansion of Greek exports of manufactured products.

Thus, the expansion of markets due to the Greek membership of the E.E.C. (E.E.C. provides about 41 percent of the country's imports and absorbs about 50 percent of its exports) and the positive results entailed, are surely going to give new dimensions to the Greek economy and will contribute to accelerate the growth of Greek exports.

Concluding, we can say that it is not out of reach, to try to increase domestic production and, hence exports, with prospects of an improved trade balance.

APPENDIX

TIME SERIES

In this appendix we present the final series of data used for the estimation of our equations. For the sources and the construction of these series see section 1 of chapters V and VI respectively, as well as the relevant sections of the same chapters.

TABLE A.1

INDEX OF VOLUME OF IMPORTED FOOD (SITC: Section 0)
(1970 = 100)

	JAN	FEB	MARCH	APRIL	MAY	JUNE	JULY	AUG	SEPT	OCT	NOV	DEC
1954	33.9	26.9	29.4	48.1	46.0	43.2	35.9	26.4	20.6	18.9	24.4	34.8
1955	36.2	34.2	32.3	40.0	65.9	102.3	33.4	25.3	29.5	34.6	45.7	67.2
1956	46.5	81.1	72.3	64.7	57.6	51.7	44.7	36.1	34.9	58.9	57.2	86.5
1957	55.2	45.9	36.5	111.6	76.0	47.2	95.4	43.3	37.9	35.5	38.3	55.2
1958	52.3	47.2	91.7	76.8	64.5	39.5	40.1	51.6	42.3	67.1	56.1	57.2
1959	52.0	45.3	63.5	65.8	40.4	47.4	39.0	28.9	46.3	65.2	53.1	57.7
1960	44.2	44.9	42.8	61.2	49.0	43.5	37.1	41.6	45.6	45.1	68.3	69.9
1961	61.4	45.2	53.1	66.0	72.7	55.3	57.1	57.5	63.2	69.8	65.8	72.3
1962	62.3	44.0	62.9	53.9	50.8	41.9	82.5	49.7	54.7	67.8	59.9	75.2
1963	50.7	69.0	89.2	70.6	56.0	81.2	70.7	62.5	54.3	83.6	79.1	100.2
1964	66.0	65.2	72.0	83.1	50.1	62.0	80.1	70.7	87.6	78.6	78.8	87.2
1965	71.4	73.0	75.6	65.4	101.1	168.8	92.0	65.7	99.5	96.1	135.3	88.8
1966	123.6	67.7	92.8	58.8	93.9	122.8	110.9	117.8	93.6	95.8	107.4	85.1
1967	79.3	90.0	130.8	98.2	120.1	85.9	99.7	82.9	108.5	77.8	116.9	140.9
1968	96.8	111.9	95.4	71.9	123.6	93.9	84.4	87.1	76.6	96.9	149.6	144.7
1969	113.3	78.1	138.2	97.6	197.8	104.6	105.1	90.6	113.1	102.1	114.8	125.4
1970	102.4	114.1	108.5	107.1	81.4	87.1	98.7	87.7	100.7	110.3	98.7	112.6
1971	75.1	58.7	85.0	102.8	106.2	98.6	112.1	98.4	103.8	112.6	248.4	230.8
1972	78.3	96.0	90.0	108.5	119.0	109.3	94.8	88.4	104.1	131.4	86.3	111.3
1973	105.2	79.2	125.2	102.9	131.5	137.4	123.5	128.0	134.4	290.9	113.4	93.5
1974	108.3	126.8	95.0	67.1	66.2	268.4	53.1	70.2	64.5	101.1	64.2	376.1
1975	99.1	125.3	297.0	106.7	55.5	123.0	92.7	78.5	102.0	100.9	61.5	149.5
1976	108.5	97.5	133.3	119.1	123.9	71.0	110.7	110.5	107.2	100.2	133.1	241.0

TABLE A.2

INDEX OF VOLUME OF IMPORTED RAW MATERIALS (SITC: Section 2)
(1970 = 100)

	JAN	FEB	MARCH	APRIL	MAY	JUNE	JULY	AUG	SEPT	OCT	NOV	DEC
1954	23.1	29.5	19.6	24.0	24.4	25.3	21.7	23.5	26.9	22.8	27.6	28.2
1955	22.4	22.1	29.8	24.8	24.7	26.1	26.2	27.4	20.7	19.7	24.2	24.9
1956	25.0	24.0	21.4	26.6	23.5	25.7	26.3	29.6	23.8	29.2	28.9	21.7
1957	28.4	26.2	23.4	27.5	33.0	29.7	40.2	40.8	39.1	28.9	29.4	35.2
1958	32.7	28.3	29.9	30.6	28.6	27.1	32.5	29.5	29.4	35.6	35.5	35.5
1959	29.8	22.2	22.6	27.6	27.8	28.1	26.9	28.1	28.3	28.1	26.8	29.8
1960	28.2	25.8	37.1	26.4	37.5	39.0	38.3	38.8	36.0	30.4	31.2	41.4
1961	29.8	41.6	32.5	35.0	36.3	35.2	36.7	30.4	31.2	27.4	38.1	28.8
1962	35.4	31.7	42.4	27.2	31.8	33.5	35.1	32.4	29.4	46.2	31.8	41.0
1963	37.6	39.2	33.7	45.4	34.9	40.6	44.8	43.2	53.0	58.2	44.6	53.9
1964	36.1	34.4	41.5	41.6	25.4	52.8	51.5	50.5	50.5	49.6	55.1	53.3
1965	50.3	54.1	57.8	58.8	56.1	63.1	50.1	49.7	71.9	61.8	58.8	63.4
1966	72.9	50.3	72.8	45.5	62.9	75.1	66.1	58.7	72.0	63.1	79.0	75.2
1967	47.7	64.9	72.1	51.0	67.1	56.2	64.2	64.3	58.2	54.4	75.5	74.4
1968	59.7	71.8	55.9	57.6	82.5	67.1	74.0	59.9	72.9	79.2	90.9	96.2
1969	85.3	75.9	102.7	67.6	92.1	81.1	88.2	77.1	80.6	92.8	84.0	92.2
1970	62.8	75.7	80.9	89.7	89.0	95.5	101.9	73.4	98.3	84.2	98.1	93.1
1971	80.4	76.6	82.9	80.3	140.4	114.0	102.8	77.5	101.3	94.9	114.3	125.6
1972	88.8	90.4	86.5	103.3	154.9	137.3	106.1	104.6	146.0	108.1	127.5	130.8
1973	154.1	130.7	143.7	131.9	157.1	117.0	189.7	151.4	142.3	138.6	128.0	152.8
1974	118.5	107.9	140.2	126.3	129.3	129.3	112.4	104.8	102.4	201.1	68.2	128.5
1975	104.5	115.0	101.0	110.6	137.8	150.6	116.6	97.1	151.2	154.4	103.2	191.1
1976	144.5	114.9	122.2	144.4	148.4	106.1	164.5	120.8	128.8	133.3	177.8	156.8

TABLE A.3

INDEX OF VOLUME OF IMPORTED FUELS (SITC: Section 3)
(1970 = 100)

	JAN	FEB	MARCH	APRIL	MAY	JUNE	JULY	AUG	SEPT	OCT	NOV	DEC
1954	31.4	23.6	29.3	24.2	27.7	27.9	33.0	28.9	34.3	30.6	34.5	32.8
1955	32.7	27.8	30.7	31.7	31.0	30.7	26.5	33.4	38.3	33.7	32.3	37.4
1956	29.9	22.6	27.1	28.9	34.0	29.5	38.8	29.8	31.6	35.4	33.5	28.7
1957	34.2	27.1	26.8	35.9	29.3	25.9	37.4	33.1	28.7	36.6	37.2	35.0
1958	38.8	28.1	39.2	32.3	34.7	35.1	40.5	31.6	35.4	49.7	32.7	41.9
1959	21.2	15.1	9.6	8.2	109.2	32.6	24.5	27.1	19.6	33.6	42.2	25.3
1960	32.2	29.3	30.7	63.9	44.0	31.6	19.0	30.5	56.9	45.7	37.2	42.7
1961	59.4	32.3	54.0	51.8	38.8	27.2	38.5	11.1	52.2	49.1	43.8	25.9
1962	71.2	18.1	21.4	34.1	13.0	24.9	18.1	15.6	45.7	21.0	19.2	163.6
1963	21.7	20.6	16.0	15.5	92.0	47.9	22.4	21.1	85.4	51.6	87.8	115.4
1964	17.5	31.2	24.8	35.9	22.9	64.5	25.7	21.6	19.4	32.3	24.0	199.9
1965	24.1	33.5	152.1	37.6	68.4	81.3	25.8	24.8	53.4	120.1	34.1	168.4
1966	45.3	29.0	59.3	30.1	59.5	72.8	33.4	27.4	77.7	38.8	49.2	298.4
1967	24.2	31.0	56.2	52.7	66.3	52.0	34.1	51.4	34.6	39.7	75.8	391.0
1968	27.9	24.6	14.9	60.3	36.4	21.8	69.7	34.8	26.3	148.7	175.0	246.1
1969	40.7	19.7	174.9	52.0	60.5	120.8	74.2	44.9	25.1	104.1	48.3	270.2
1970	54.3	6.7	90.1	149.3	68.1	82.2	154.3	65.7	120.9	123.0	124.2	168.1
1971	82.5	75.1	73.9	96.2	64.2	62.6	54.1	83.0	78.0	93.5	154.4	371.3
1972	92.1	96.6	64.2	64.5	124.0	103.5	40.6	107.5	84.8	371.1	237.2	345.2
1973	128.6	141.1	140.1	84.9	265.6	158.4	187.9	163.7	167.6	281.5	208.4	439.7
1974	149.1	164.8	181.1	131.3	298.9	194.2	212.7	125.3	149.1	196.6	226.5	171.0
1975	131.2	163.2	254.7	211.5	209.0	188.4	126.7	96.8	183.1	202.6	228.3	258.5
1976	96.9	137.6	121.1	122.5	185.5	189.7	135.8	131.8	277.7	255.6	162.6	333.1

TABLE A.4

INDEX OF VOLUME OF IMPORTED CHEMICALS (SITC: Section 5)
(1970 = 100)

	JAN	FEB	MARCH	APRIL	MAY	JUNE	JULY	AUG	SEPT	OCT	NOV	DEC
1954	18.9	16.8	10.6	10.4	6.6	8.0	13.5	12.1	18.8	12.4	27.7	13.5
1955	11.0	14.5	9.9	16.9	11.9	8.7	14.6	15.0	16.4	16.8	17.0	11.7
1956	12.1	16.8	10.4	10.5	13.3	9.5	11.3	11.3	13.9	31.6	23.5	23.6
1957	17.9	16.6	16.0	22.0	23.4	14.8	17.5	16.0	26.6	25.7	29.1	17.9
1958	34.0	37.9	31.2	23.7	26.8	20.3	30.7	14.6	26.9	46.6	28.5	36.3
1959	33.9	15.8	18.0	25.4	16.4	18.2	26.5	32.9	32.2	37.7	28.7	38.0
1960	34.8	26.9	33.0	21.6	24.5	21.9	25.9	38.0	24.4	32.6	36.3	33.1
1961	31.9	24.1	30.6	26.3	24.9	29.4	34.9	32.1	40.7	41.0	35.0	27.9
1962	34.8	31.0	36.9	28.7	41.6	37.9	32.7	43.5	48.1	50.7	42.8	42.4
1963	49.1	43.3	38.4	42.9	42.4	43.4	51.1	58.3	49.5	68.0	67.6	51.4
1964	35.5	59.0	52.7	41.2	44.6	50.7	58.0	57.2	52.4	58.2	63.5	56.1
1965	54.3	56.4	54.8	48.6	63.0	51.7	54.9	67.9	52.4	63.8	54.0	62.8
1966	74.4	57.0	63.3	46.8	66.4	62.1	65.8	51.9	74.9	59.3	75.0	64.7
1967	44.6	68.0	68.5	65.1	72.3	59.9	56.4	52.6	52.1	56.1	81.6	92.2
1968	63.2	69.9	60.2	64.7	77.1	63.0	62.1	48.7	62.7	86.9	80.1	74.4
1969	72.8	59.0	87.1	75.9	94.9	80.1	75.3	75.9	81.6	79.5	88.8	91.7
1970	98.3	81.0	91.2	79.6	96.4	114.6	93.4	75.8	88.0	92.7	98.0	117.8
1971	107.6	98.6	85.8	92.5	114.1	107.9	117.6	77.6	107.8	124.5	169.4	139.1
1972	99.0	95.5	114.5	120.3	128.6	117.3	120.7	112.3	127.2	122.1	131.1	160.7
1973	140.1	125.3	157.7	139.4	150.1	169.0	162.8	143.5	168.0	167.5	158.1	132.7
1974	146.6	118.5	143.7	131.8	141.3	126.4	133.5	110.8	114.3	133.1	80.6	85.8
1975	127.1	125.8	119.1	147.6	120.6	152.9	122.1	118.1	148.4	146.2	166.7	162.6
1976	164.7	144.6	168.5	157.4	155.4	97.2	149.7	128.0	196.2	163.1	179.3	170.2

TABLE A.5

INDEX OF VOLUME OF IMPORTED MANUFACTURES (SITC: Section 6)
(1970 = 100)

	JAN	FEB	MARCH	APRIL	MAY	JUNE	JULY	AUG	SEPT	OCT	NOV	DEC
1954	26.0	23.9	27.8	30.3	30.5	26.8	29.9	34.2	40.5	36.1	42.6	33.7
1955	31.7	29.9	36.2	30.7	34.1	35.3	30.6	27.4	30.0	32.3	35.8	32.8
1956	28.7	30.4	31.8	37.3	33.1	31.9	40.3	39.5	36.2	50.3	41.5	34.1
1957	41.5	40.8	37.8	42.2	47.5	40.9	41.2	42.7	43.7	36.8	49.0	55.9
1958	41.9	42.8	49.2	50.7	58.4	44.0	48.1	44.7	50.4	54.4	51.0	49.9
1959	46.1	45.7	44.8	53.2	42.5	44.2	54.9	46.3	51.7	49.0	45.6	42.3
1960	42.2	42.9	55.5	47.7	56.7	48.9	53.6	53.7	53.5	56.5	58.8	60.1
1961	58.7	48.9	67.4	51.3	62.1	57.2	56.9	50.4	55.2	60.9	64.7	47.7
1962	60.7	61.4	71.0	62.2	66.1	63.7	62.9	69.9	64.0	66.6	71.2	62.9
1963	62.7	59.7	59.4	70.4	55.1	72.4	69.5	88.3	79.8	80.1	70.5	75.0
1964	55.2	68.6	71.5	68.2	77.2	91.5	100.7	85.2	82.7	81.0	100.3	101.3
1965	72.6	98.1	93.4	105.0	108.2	89.9	88.4	103.4	109.5	97.6	110.6	99.6
1966	99.9	75.1	114.7	74.8	102.3	111.6	114.1	96.8	59.2	107.7	123.9	102.6
1967	70.4	95.7	97.9	94.1	105.8	101.9	76.2	117.6	100.2	104.3	120.6	119.0
1968	82.6	108.4	92.4	87.1	107.5	92.8	89.4	79.4	119.4	115.2	98.4	117.0
1969	110.2	90.6	103.6	106.9	115.0	101.3	109.1	90.9	108.5	107.9	92.4	95.3
1970	84.8	88.0	117.0	103.2	130.5	103.7	94.8	95.7	114.6	102.9	128.7	104.7
1971	103.2	89.2	95.1	91.7	139.1	127.5	116.4	98.1	108.5	138.2	167.4	141.8
1972	127.5	92.4	121.5	120.6	150.8	141.7	127.6	124.3	154.8	128.5	142.7	167.2
1973	136.7	158.5	153.2	141.3	162.4	130.4	179.8	159.4	149.8	179.8	157.4	238.0
1974	194.6	149.7	201.7	137.1	182.7	175.4	129.9	127.7	147.6	150.1	117.8	122.6
1975	137.7	112.5	103.8	189.5	134.8	160.0	122.4	112.7	165.6	135.5	152.0	142.6
1976	140.3	132.7	142.9	135.8	150.6	160.7	158.7	138.7	161.9	163.7	144.6	178.5

TABLE A.6

INDEX OF VOLUME OF IMPORTED MACHINERY AND TRANSPORT EQUIPMENT (Without Ships) (SITC: Section 7) (1970 = 100)

	JAN	FEB	MARCH	APRIL	MAY	JUNE	JULY	AUG	SEPT	OCT	NOV	DEC
1954	13.5	12.0	11.8	13.9	14.3	13.7	14.2	13.4	14.9	13.3	16.4	15.0
1955	14.7	13.7	14.8	15.1	17.9	20.1	14.8	14.8	14.5	15.2	16.6	18.1
1956	15.0	17.6	16.4	17.6	17.8	18.0	17.7	16.4	15.2	23.4	20.1	19.5
1957	17.8	17.5	15.5	24.5	23.1	18.8	23.7	19.7	19.7	17.9	20.1	22.5
1958	21.4	20.1	25.8	23.7	24.8	19.3	22.1	21.6	20.7	27.2	23.9	24.9
1959	20.6	16.4	18.1	21.5	23.8	19.3	19.9	18.2	20.1	22.6	22.3	21.6
1960	19.4	19.2	20.1	21.9	28.1	22.6	19.3	22.7	23.2	24.2	25.6	27.3
1961	26.4	21.6	27.1	26.3	28.2	26.7	28.0	22.8	26.5	29.2	28.9	23.6
1962	31.7	32.3	37.3	31.5	34.6	30.1	28.4	30.3	29.1	35.6	34.7	30.4
1963	27.7	27.1	28.6	37.9	34.9	36.5	35.6	47.0	36.2	38.6	40.9	43.5
1964	26.7	40.4	39.8	36.3	51.2	52.4	57.4	55.6	42.0	41.7	49.7	48.8
1965	43.6	48.0	50.0	59.1	69.1	57.6	45.7	52.1	56.5	47.5	51.2	59.8
1966	44.3	36.6	46.1	49.9	64.8	66.5	47.8	52.4	52.1	49.5	68.7	60.5
1967	43.9	49.1	55.9	53.8	70.0	63.1	46.8	67.2	63.1	61.2	56.3	60.2
1968	47.4	52.1	54.9	63.7	83.9	94.8	70.9	62.9	61.3	99.7	62.7	97.5
1969	95.9	102.3	66.3	71.7	90.2	81.3	83.2	77.8	69.1	76.1	82.1	101.6
1970	73.5	71.1	108.0	90.8	86.0	93.5	80.1	77.0	92.3	78.9	102.0	100.4
1971	76.1	73.9	75.5	98.2	107.0	96.7	104.0	71.0	89.8	94.5	99.0	106.5
1972	79.9	84.9	111.4	125.4	136.0	114.1	118.4	117.7	123.7	113.2	128.9	140.9
1973	126.6	102.0	148.3	169.5	135.5	248.0	156.2	132.5	177.8	149.0	152.8	160.8
1974	157.4	149.5	162.4	143.9	168.7	164.1	139.6	124.7	119.1	150.5	106.4	108.9
1975	154.0	141.5	145.8	125.8	123.7	141.4	124.0	119.7	140.5	145.0	125.2	145.7
1976	141.1	117.2	124.3	138.7	151.4	94.8	167.5	146.9	160.9	188.6	172.7	180.1

TABLE A.7

INDEX OF VOLUME OF TOTAL IMPORTS (SITC: Sections 0-9)
(1970 = 100)

	JAN	FEB	MARCH	APRIL	MAY	JUNE	JULY	AUG	SEPT	OCT	NOV	DEC
1954	19.1	17.1	16.8	19.8	20.4	19.5	20.2	19.0	21.1	18.9	23.2	21.3
1955	20.8	19.5	21.0	21.4	25.4	28.6	21.0	21.0	20.6	21.6	23.6	25.7
1956	21.3	25.0	23.2	25.0	25.2	25.6	25.1	23.2	21.6	33.3	28.5	27.7
1957	25.2	24.9	22.0	34.7	32.8	26.6	33.7	28.0	28.0	25.4	28.6	31.9
1958	30.3	28.5	36.6	33.7	35.2	27.5	31.4	30.6	29.4	38.6	33.9	35.4
1959	29.2	23.4	25.7	30.5	33.8	27.5	28.2	25.8	28.6	32.1	31.7	30.7
1960	27.6	27.3	28.6	31.1	39.9	32.0	27.4	32.2	33.0	34.4	36.4	38.8
1961	37.5	30.7	38.6	37.3	40.0	37.9	39.8	32.4	37.6	41.5	41.1	33.6
1962	43.1	35.8	43.6	37.2	40.2	37.5	39.2	38.5	40.5	45.2	41.7	55.0
1963	39.1	39.0	39.9	45.3	45.8	48.2	47.2	52.6	53.0	57.1	56.9	61.8
1964	36.4	47.0	47.9	46.2	47.2	59.8	62.4	57.1	52.6	53.5	60.2	75.6
1965	51.9	57.9	71.1	62.3	74.7	73.9	56.0	60.1	69.3	73.6	68.2	79.5
1966	67.9	49.4	68.1	51.4	72.2	79.2	66.9	63.2	71.5	64.7	80.8	93.5
1967	50.2	62.6	72.7	64.9	79.9	68.4	58.5	71.5	67.5	65.8	79.7	113.8
1968	59.1	68.4	61.3	66.2	85.3	78.1	73.0	62.2	69.8	100.7	92.5	114.2
1969	88.2	78.9	95.7	77.3	101.5	88.6	86.7	76.4	77.6	88.0	84.7	114.5
1970	77.5	74.5	100.1	97.2	91.9	95.2	94.5	78.2	98.2	92.0	110.6	111.8
1971	84.7	78.3	80.6	93.1	112.1	101.0	102.5	81.2	96.8	105.4	139.4	154.9
1972	91.0	89.1	102.7	111.3	135.5	118.9	108.2	110.9	125.2	139.7	135.7	160.3
1973	129.6	119.5	145.4	140.7	155.7	176.4	163.2	142.8	159.8	184.7	151.5	188.4
1974	153.3	140.7	161.4	129.5	163.9	173.8	129.9	114.7	118.6	150.0	104.8	147.4
1975	132.7	130.1	156.7	146.2	125.4	149.3	118.6	108.6	145.9	143.6	134.9	162.2
1976	136.8	123.0	139.2	138.2	151.3	114.6	153.7	134.9	164.5	169.3	163.2	198.1

TABLE A.8

RELATIVE IMPORT PRICES OF FOOD

	JAN	FEB	MARCH	APRIL	MAY	JUNE	JULY	AUG	SEPT	OCT	NOV	DEC
1954	145.8	142.6	143.7	149.2	151.0	148.9	143.6	143.8	150.1	158.6	139.3	131.5
1955	135.8	153.1	151.7	149.5	148.1	142.8	139.8	133.5	131.8	137.1	135.0	123.1
1956	124.0	128.9	125.3	132.3	123.1	122.9	124.3	127.5	127.2	127.4	117.2	124.0
1957	120.4	113.5	119.0	124.2	127.7	124.4	138.3	129.1	130.4	128.2	134.4	127.3
1958	127.1	132.4	127.9	121.6	126.7	120.8	119.3	112.7	116.7	105.9	112.7	123.1
1959	120.4	115.4	109.7	107.4	123.6	112.2	121.9	118.2	109.4	128.2	111.3	112.9
1960	113.0	112.6	118.3	108.5	115.1	122.3	122.5	122.2	111.1	113.6	93.4	103.3
1961	94.7	104.9	107.9	96.7	91.8	110.1	110.3	104.1	103.7	99.0	108.8	99.3
1962	103.3	115.6	108.0	111.0	113.8	120.5	104.0	104.1	105.1	90.2	92.7	91.7
1963	99.4	84.4	81.6	94.8	84.6	111.4	94.3	107.0	102.9	85.2	89.5	94.0
1964	86.9	90.9	123.9	102.0	111.1	112.5	107.4	100.8	91.3	96.0	99.7	100.6
1965	97.0	98.1	104.5	96.1	92.4	82.7	99.4	98.8	87.3	90.8	84.9	91.9
1966	85.2	88.3	91.0	95.2	87.6	84.9	93.3	87.6	94.4	86.0	89.5	84.1
1967	88.4	88.3	78.0	86.3	92.1	99.6	92.2	105.4	89.0	95.9	95.7	92.5
1968	94.9	89.5	91.5	93.2	94.6	97.3	101.8	94.6	90.7	87.7	80.0	80.8
1969	82.4	82.3	76.7	83.2	75.9	87.5	89.4	90.5	89.5	90.2	83.8	91.6
1970	91.0	84.5	89.8	77.2	115.6	97.0	104.0	99.7	95.8	87.5	91.8	91.1
1971	93.9	94.2	93.9	99.7	87.5	98.9	97.6	97.5	97.7	96.4	78.7	80.4
1972	99.8	102.9	100.2	95.8	112.2	118.5	114.6	113.6	108.4	86.1	102.1	102.9
1973	103.4	105.5	107.0	116.8	122.2	119.9	110.3	108.2	103.5	88.1	90.8	92.0
1974	92.6	100.9	98.4	102.8	98.9	118.0	113.2	113.6	103.0	105.2	99.4	113.3
1975	97.2	98.5	103.2	95.5	100.4	117.8	113.0	115.6	115.8	161.2	108.6	135.3
1976	107.3	96.2	109.6	105.0	100.0	105.0	101.2	103.5	102.0	104.0	101.4	120.4

TABLE A.9

RELATIVE IMPORT PRICES OF RAW MATERIALS

	JAN	FEB	MARCH	APRIL	MAY	JUNE	JULY	AUG	SEPT	OCT	NOV	DEC
1954	137.7	128.0	139.3	134.8	135.1	131.5	133.8	134.1	138.5	130.5	132.6	128.9
1955	131.8	127.2	109.2	111.5	126.6	121.3	123.3	133.0	137.5	123.5	126.9	130.0
1956	130.2	127.3	128.6	125.5	133.0	123.4	133.0	136.0	136.8	132.5	137.2	133.9
1957	128.1	126.5	130.6	128.0	124.7	125.8	127.1	130.6	132.5	135.8	132.5	130.8
1958	133.2	125.6	128.8	125.4	129.3	130.4	128.7	131.4	130.8	127.9	128.2	126.8
1959	124.5	133.7	130.2	127.4	123.8	122.7	123.9	131.2	131.6	130.4	130.4	123.3
1960	114.9	119.1	110.8	123.3	125.0	117.9	120.4	127.1	126.6	123.3	125.2	126.7
1961	122.2	113.4	116.2	115.9	114.1	121.3	121.7	113.3	135.7	130.8	129.3	128.6
1962	125.7	124.9	124.1	126.4	125.4	121.9	128.0	124.4	128.7	122.6	121.1	119.3
1963	116.5	122.2	121.4	120.4	124.0	125.6	115.9	127.8	127.2	130.1	125.5	130.5
1964	123.4	125.5	122.9	124.9	122.6	127.6	124.1	126.9	127.0	123.3	115.1	113.8
1965	119.0	115.4	116.5	119.1	121.1	115.7	118.8	114.8	119.5	119.7	118.4	118.0
1966	117.3	116.7	120.2	114.9	116.0	113.4	120.9	113.2	118.0	121.6	120.8	113.7
1967	114.3	110.8	112.6	115.3	110.8	111.7	112.8	115.4	114.5	115.6	113.8	105.3
1968	111.2	104.4	104.9	114.4	111.1	108.5	107.8	108.2	107.8	111.6	110.6	114.0
1969	103.0	101.7	107.9	103.0	99.3	103.6	102.3	107.6	104.5	103.8	106.8	109.8
1970	102.3	102.7	103.7	109.0	104.7	106.5	113.3	113.9	118.4	109.4	106.3	111.6
1971	103.2	103.1	101.4	91.8	106.4	101.3	102.8	100.8	100.0	101.1	93.8	96.1
1972	98.5	92.6	96.1	99.3	94.9	95.5	87.8	95.9	93.8	82.2	87.1	88.6
1973	84.7	81.2	95.5	79.9	84.2	82.3	89.4	89.9	83.9	78.1	73.6	69.2
1974	76.1	69.7	75.4	81.3	87.4	85.1	85.5	88.5	90.2	97.4	104.3	100.6
1975	102.1	99.4	107.7	104.0	131.2	88.5	91.1	93.1	101.2	93.2	93.2	87.5
1976	94.1	83.1	83.6	85.3	83.5	87.7	85.9	85.2	86.5	84.1	83.3	91.4

TABLE A.10

RELATIVE IMPORT PRICES OF FUELS

	JAN	FEB	MARCH	APRIL	MAY	JUNE	JULY	AUG	SEPT	OCT	NOV	DEC
1954	210.5	212.6	210.1	211.0	206.7	208.4	209.6	206.5	198.6	198.3	195.2	199.8
1955	204.6	207.3	214.6	202.0	204.2	190.1	200.5	201.6	202.0	206.5	211.8	203.7
1956	200.2	201.5	192.6	202.8	194.1	193.7	208.2	196.9	209.2	205.6	199.7	216.4
1957	242.2	254.1	260.5	265.7	255.5	260.8	237.6	230.9	227.6	215.3	219.8	223.1
1958	207.0	210.7	201.5	205.7	207.9	214.1	191.9	194.9	143.0	183.0	210.7	181.7
1959	200.5	170.8	169.9	177.0	134.6	123.6	139.4	138.1	144.0	127.2	130.2	129.4
1960	125.6	140.9	108.1	102.2	109.0	117.1	114.1	112.8	98.6	105.1	102.0	109.7
1961	100.0	93.4	94.2	95.8	105.4	100.7	106.4	112.5	102.1	102.7	115.6	126.9
1962	111.1	161.6	127.4	108.6	130.6	112.0	135.9	133.6	101.4	104.2	126.3	90.3
1963	123.4	111.1	115.7	126.8	108.1	103.8	122.9	134.3	108.2	104.2	107.1	103.0
1964	109.8	123.9	115.0	109.6	127.8	106.7	117.5	124.2	129.0	108.6	133.7	98.4
1965	124.8	115.5	100.8	103.0	99.0	96.5	120.2	120.7	98.2	98.8	114.9	92.6
1966	124.3	124.2	104.8	104.2	103.7	94.4	107.8	109.8	96.4	107.2	105.4	77.0
1967	124.5	113.5	97.7	92.0	94.3	95.9	109.5	101.7	111.3	109.8	106.3	78.0
1968	129.2	137.2	141.4	95.1	101.1	114.3	97.0	95.7	117.6	93.4	94.3	93.9
1969	116.3	131.0	97.1	102.2	96.1	98.2	104.0	89.3	114.8	94.0	113.4	89.4
1970	107.7	103.7	89.5	88.2	85.5	94.0	93.3	91.5	89.6	97.6	87.8	93.1
1971	110.1	104.5	96.8	101.1	99.7	105.9	88.2	106.3	106.9	104.9	87.9	93.7
1972	98.0	110.6	100.9	99.6	97.7	101.6	138.8	119.6	94.3	107.7	104.4	111.8
1973	110.7	113.1	122.4	103.1	106.7	118.7	117.0	102.9	98.2	93.5	100.3	101.4
1974	109.0	127.0	112.6	190.7	221.4	210.5	179.4	205.6	183.9	229.2	242.1	212.0
1975	228.8	214.3	167.8	215.2	234.2	204.0	210.0	214.2	244.8	224.1	254.4	256.0
1976	271.7	238.2	232.5	231.7	242.4	236.1	241.8	217.4	255.2	241.3	242.1	262.8

TABLE A.11

RELATIVE IMPORT PRICES OF CHEMICALS

	JAN	FEB	MARCH	APRIL	MAY	JUNE	JULY	AUG	SEPT	OCT	NOV	DEC
1954	161.2	147.5	169.7	160.5	179.1	166.6	180.3	181.0	160.1	164.7	172.0	164.3
1955	178.1	154.2	172.4	185.4	171.9	162.6	166.0	163.5	173.1	190.2	167.5	170.5
1956	185.5	210.0	154.2	176.3	148.2	170.1	145.2	161.7	152.8	131.5	124.4	124.2
1957	128.7	122.1	122.3	125.9	135.8	136.5	131.1	138.5	143.6	137.4	137.7	165.7
1958	144.1	112.7	98.0	107.4	121.9	135.8	119.3	150.8	131.4	118.7	129.9	118.7
1959	119.8	157.8	151.3	126.6	153.1	146.4	132.3	125.8	121.6	128.6	119.9	120.4
1960	125.4	113.0	122.7	130.6	144.9	154.4	135.2	120.6	138.1	131.4	112.9	118.5
1961	123.3	131.6	141.8	130.6	141.5	134.8	132.0	118.9	122.3	132.8	133.4	132.9
1962	129.6	130.9	122.1	130.1	102.3	121.2	130.7	126.1	121.2	116.7	113.8	110.8
1963	102.6	103.4	112.5	110.6	113.9	117.7	103.8	110.1	113.5	96.8	99.3	113.6
1964	112.8	106.7	115.6	109.8	107.0	121.0	113.0	106.8	100.8	104.8	99.1	103.9
1965	107.2	112.1	105.7	113.6	108.1	112.3	110.6	107.8	113.1	106.8	110.4	113.0
1966	104.2	117.3	95.8	122.2	115.3	114.3	104.7	110.9	107.8	111.2	106.9	115.4
1967	104.2	106.2	114.7	112.1	116.3	117.7	116.5	125.6	123.7	125.2	117.8	113.4
1968	110.6	109.1	116.2	117.5	113.2	137.4	127.4	125.3	123.0	106.5	101.8	106.2
1969	107.9	110.6	117.6	112.6	103.0	118.5	140.3	104.7	107.2	108.7	105.4	95.8
1970	96.1	110.2	108.8	111.4	109.5	82.7	97.1	102.4	104.9	104.2	105.4	88.6
1971	89.2	97.9	107.2	101.4	107.4	106.9	94.1	108.3	106.8	85.7	81.9	87.6
1972	112.0	106.6	103.5	108.1	114.2	115.5	111.9	103.9	96.8	98.6	102.5	86.1
1973	93.8	97.0	95.6	113.3	92.9	99.4	98.7	92.7	91.7	86.3	79.7	82.9
1974	82.6	90.8	94.7	93.5	102.2	112.0	112.3	111.5	106.5	108.2	113.1	106.2
1975	109.9	112.4	104.7	113.7	116.1	111.8	108.1	109.1	104.9	103.1	97.6	98.6
1976	101.0	106.0	101.9	107.4	106.2	113.0	108.2	104.9	89.3	101.1	102.8	103.3

TABLE A.12

RELATIVE IMPORT PRICES OF MANUFACTURES

	JAN	FEB	MARCH	APRIL	MAY	JUNE	JULY	AUG	SEPT	OCT	NOV	DEC
1954	109.3	106.7	104.5	105.3	106.0	101.2	99.8	98.8	94.6	98.8	100.2	101.5
1955	100.4	97.0	100.7	99.9	97.9	97.9	94.9	98.2	98.9	96.7	99.4	98.8
1956	98.1	98.5	95.1	96.4	96.9	101.7	102.3	102.2	94.8	98.1	96.7	100.8
1957	92.8	93.5	91.2	94.8	96.5	96.6	97.0	96.7	92.7	94.3	93.3	96.0
1958	94.8	93.4	94.3	89.0	87.4	88.9	87.8	85.0	101.4	88.2	85.1	84.9
1959	87.2	85.4	85.5	87.5	87.0	88.3	82.8	86.8	88.8	90.6	93.4	95.3
1960	89.3	89.7	85.2	92.5	91.6	89.0	83.1	89.9	86.7	87.7	86.4	87.1
1961	83.3	84.5	85.7	88.2	86.5	88.3	86.5	83.8	88.1	86.7	84.7	83.7
1962	81.4	82.5	85.3	81.2	82.5	82.1	82.0	81.7	80.9	80.3	76.9	77.4
1963	78.8	79.5	83.1	79.0	80.4	75.7	76.5	77.5	78.6	78.6	78.8	70.7
1964	77.0	76.7	80.9	76.1	78.7	74.6	77.5	79.1	79.2	80.9	73.2	69.5
1965	83.2	70.5	78.2	70.5	76.3	78.2	76.3	73.2	79.2	80.4	76.4	72.4
1966	73.5	76.8	73.4	79.8	77.9	75.4	70.2	71.8	78.8	73.2	75.6	76.7
1967	78.7	76.4	77.6	78.1	72.4	72.0	78.0	73.1	76.5	77.7	76.3	77.1
1968	82.5	78.9	79.0	77.8	78.9	79.4	80.9	80.6	81.2	80.1	80.7	77.2
1969	79.3	83.2	85.4	86.5	83.7	85.1	81.0	86.4	87.4	85.8	87.9	87.7
1970	93.1	89.1	88.7	87.7	89.2	89.3	86.7	87.5	86.9	85.7	87.9	85.6
1971	85.5	85.7	85.9	86.9	86.1	77.4	81.2	84.8	98.7	72.3	80.4	79.9
1972	82.4	88.7	82.6	86.3	88.5	91.8	85.9	92.9	85.3	86.7	97.0	83.2
1973	92.7	88.3	96.2	97.9	92.8	93.2	89.2	90.8	88.7	78.9	78.2	83.1
1974	81.5	81.8	81.2	83.9	89.6	92.2	89.6	93.8	97.5	97.2	91.0	94.7
1975	104.5	99.0	86.8	90.9	88.0	92.6	86.7	89.9	90.0	84.8	86.7	85.8
1976	88.9	87.9	82.2	86.3	85.3	88.4	80.5	84.3	84.0	85.0	87.9	86.0

TABLE A.13

RELATIVE IMPORT PRICES OF MACHINERY
AND TRANSPORT EQUIPMENT (Without Ships)

	JAN	FEB	MARCH	APRIL	MAY	JUNE	JULY	AUG	SEPT	OCT	NOV	DEC
1954	143.9	153.2	150.8	153.0	128.2	144.3	147.3	151.7	155.1	143.8	140.2	141.7
1955	122.6	127.2	142.0	128.9	117.7	127.5	128.6	121.3	119.0	116.8	126.0	108.3
1956	130.4	125.9	134.1	130.0	118.2	127.1	133.5	125.0	126.3	142.6	132.9	124.6
1957	130.9	115.4	128.2	130.1	128.9	129.7	128.9	142.3	136.7	137.3	136.5	141.4
1958	132.2	141.6	141.0	132.9	135.8	137.3	125.9	129.2	129.2	122.3	126.6	121.5
1959	133.1	142.2	130.0	140.3	141.0	137.5	139.5	137.9	130.5	135.2	141.3	126.9
1960	137.5	131.6	147.3	133.5	115.3	121.7	130.1	123.4	112.4	122.9	114.1	123.2
1961	110.8	117.1	118.5	118.8	122.8	121.3	124.9	121.1	128.9	124.8	123.5	120.9
1962	123.2	125.8	117.4	124.3	132.8	135.9	133.5	133.1	140.2	121.5	111.6	125.8
1963	123.9	118.4	113.6	125.0	130.1	118.3	130.5	128.3	125.3	127.0	117.1	115.5
1964	121.5	118.0	118.7	122.2	118.1	123.2	124.2	118.8	119.9	120.4	119.0	112.8
1965	116.8	106.1	117.1	110.4	105.0	112.9	113.2	118.0	116.6	121.3	120.7	170.6
1966	120.9	120.5	122.2	118.5	111.9	103.1	117.3	130.7	129.1	123.1	118.4	118.4
1967	129.4	122.7	123.6	122.3	117.9	123.9	131.0	124.1	119.8	116.7	121.9	116.0
1968	116.7	120.7	111.5	117.5	118.7	120.2	127.6	120.3	121.6	109.2	117.1	100.3
1969	102.2	108.7	116.7	110.4	109.7	114.8	107.6	107.1	108.9	109.3	102.4	97.7
1970	100.1	109.6	101.5	108.1	108.0	114.1	118.1	109.9	98.6	114.2	109.8	105.5
1971	110.3	102.3	105.5	97.9	110.3	115.8	115.1	123.7	115.8	111.6	108.7	106.8
1972	116.4	114.0	110.8	109.1	110.8	115.0	115.0	114.3	109.9	113.1	102.1	100.6
1973	103.9	111.4	105.2	91.3	97.4	97.2	102.7	98.2	95.7	85.6	82.3	77.7
1974	72.5	76.9	80.8	80.6	81.1	77.9	77.8	80.5	84.9	83.1	81.3	84.2
1975	80.3	83.0	87.2	85.2	90.0	91.1	91.1	98.8	92.0	96.6	91.8	95.9
1976	88.1	98.0	94.1	94.6	93.9	102.9	86.3	99.6	87.1	87.6	94.8	100.4

TABLE A.14

RELATIVE PRICES OF TOTAL IMPORTS

	JAN	FEB	MARCH	APRIL	MAY	JUNE	JULY	AUG	SEPT	OCT	NOV	DEC
1954	153.8	150.3	154.2	154.6	150.8	151.6	152.6	153.4	150.6	148.9	148.0	145.8
1955	141.8	145.3	148.0	145.8	141.2	141.8	140.2	139.4	141.4	141.2	142.6	133.9
1956	140.8	142.4	136.9	139.9	132.7	136.1	141.1	140.3	137.2	140.7	134.1	134.5
1957	136.6	131.2	136.4	137.2	138.0	139.3	140.2	143.6	141.2	140.9	141.1	144.2
1958	138.2	134.8	135.9	131.3	135.9	137.3	128.8	131.1	129.9	123.3	127.2	125.2
1959	127.5	132.7	127.0	126.1	129.2	125.9	126.9	124.5	122.7	128.6	126.4	122.2
1960	122.7	122.0	129.5	122.7	117.1	121.2	122.3	119.6	112.6	116.5	110.5	115.4
1961	106.7	110.4	112.6	109.9	112.8	116.7	117.5	113.4	118.0	117.9	119.8	116.7
1962	113.2	119.5	114.2	115.5	115.7	117.7	119.3	116.5	115.6	108.1	105.7	101.5
1963	107.0	103.8	105.3	108.7	107.5	110.4	108.2	112.8	109.0	105.4	103.5	103.8
1964	106.7	105.0	114.0	109.6	109.6	110.2	111.6	108.2	107.3	107.6	104.5	98.6
1965	106.5	99.8	102.6	99.9	99.9	101.7	105.9	103.5	104.4	103.4	105.5	100.0
1966	103.1	106.0	102.3	106.9	102.7	97.8	102.4	107.0	107.9	104.6	104.3	93.4
1967	108.4	103.5	102.9	104.4	103.1	106.6	110.1	109.1	107.6	107.8	107.1	94.1
1968	107.4	105.4	104.6	104.3	106.0	110.1	109.7	107.5	107.4	98.7	98.8	94.4
1969	97.1	100.4	100.8	102.5	98.1	103.2	103.1	100.9	104.3	101.4	101.4	94.7
1970	100.0	103.7	99.4	98.6	105.8	100.8	103.8	104.1	98.9	102.5	102.7	97.6
1971	99.6	97.3	100.0	97.2	101.1	102.1	101.0	107.2	107.1	96.8	93.6	93.4
1972	103.5	104.5	101.4	101.9	103.7	107.7	106.8	108.5	99.3	99.6	99.4	96.3
1973	98.0	98.9	102.9	96.7	98.6	99.0	100.6	98.5	94.9	86.5	83.3	83.8
1974	79.9	84.4	84.2	92.7	106.9	101.6	99.4	100.6	101.4	107.8	116.7	109.9
1975	104.3	104.4	102.7	106.5	113.2	106.3	103.8	106.8	110.8	113.8	114.2	117.2
1976	102.3	107.2	101.0	103.1	107.2	115.8	99.9	104.6	111.4	106.9	104.0	120.4

TABLE A.15

INDEX OF INDUSTRIAL PRODUCTION OF GREECE
(1970 = 100)

	JAN	FEB	MARCH	APRIL	MAY	JUNE	JULY	AUG	SEPT	OCT	NOV	DEC
1954	26.9	26.5	27.4	27.4	27.1	27.2	28.4	28.4	29.5	30.5	29.9	30.1
1955	28.5	27.6	29.2	28.8	30.1	30.3	30.5	29.7	29.2	29.0	29.2	28.6
1956	28.5	27.2	28.8	29.4	29.9	30.3	30.3	29.4	30.3	30.8	30.5	29.9
1957	29.6	29.0	29.7	30.8	32.3	30.8	32.8	32.5	32.8	34.7	34.7	32.7
1958	32.8	30.5	33.4	33.8	34.7	34.8	36.9	37.0	37.2	37.8	37.4	36.3
1959	35.8	37.1	38.5	40.4	38.9	39.2	38.0	39.0	41.0	41.0	41.2	38.7
1960	36.6	39.7	40.5	41.0	42.6	42.8	43.7	43.5	46.2	46.6	45.7	44.3
1961	39.2	41.8	43.8	46.0	49.9	55.6	56.4	52.8	51.9	50.5	49.0	47.2
1962	42.6	44.4	45.7	47.5	48.6	57.5	58.5	55.5	55.6	54.9	50.6	47.5
1963	43.4	46.9	48.8	48.6	51.3	59.5	62.1	62.8	63.5	60.6	54.9	54.1
1964	49.5	51.2	53.9	57.5	55.3	66.0	72.4	69.9	66.2	63.6	61.1	59.5
1965	54.5	56.0	60.8	60.5	64.2	76.9	76.7	69.6	70.7	66.5	64.7	62.3
1966	60.3	65.6	69.7	70.3	82.1	89.7	87.8	79.5	78.4	76.0	73.9	71.0
1967	68.3	70.7	74.4	73.8	82.6	88.9	86.4	82.0	77.1	78.1	78.2	74.1
1968	67.3	74.4	77.4	77.0	84.1	96.2	95.1	88.9	85.5	85.4	83.1	81.4
1969	75.2	82.1	85.4	85.7	93.1	104.0	104.6	99.0	97.9	95.6	94.5	89.6
1970	87.2	93.5	96.7	98.0	101.5	108.4	112.1	107.1	110.0	105.2	106.5	102.3
1971	96.0	103.7	106.3	104.1	109.4	119.0	125.1	121.1	118.8	114.1	116.4	114.1
1972	107.7	116.6	118.8	119.4	127.1	137.0	138.9	128.9	134.1	129.5	131.3	134.9
1973	125.8	136.9	142.2	146.4	148.1	149.9	151.7	147.6	162.2	156.0	149.6	147.7
1974	139.0	152.8	149.4	144.1	148.0	144.4	131.1	136.9	151.4	145.7	142.5	144.9
1975	134.6	144.4	152.1	152.2	145.5	151.2	148.4	145.4	161.0	154.0	159.5	158.2
1976	147.2	155.0	165.4	165.0	160.9	173.2	162.7	155.6	180.6	175.2	179.1	176.6

TABLE A.16

INDEX OF CHEMICAL PRODUCTION OF GREECE
(1970 = 100)

	JAN	FEB	MARCH	APRIL	MAY	JUNE	JULY	AUG	SEPT	OCT	NOV	DEC
1954	14.5	14.3	14.9	14.1	13.3	13.9	16.2	16.2	16.8	18.6	18.5	19.3
1955	17.5	15.7	17.4	18.3	19.0	18.2	18.9	16.9	16.5	17.0	17.4	16.2
1956	16.7	13.8	15.0	16.9	16.9	16.0	14.9	13.9	14.0	14.3	15.5	16.7
1957	16.4	15.1	15.0	16.9	18.1	15.8	18.1	19.2	18.7	21.5	22.2	19.9
1958	18.0	15.2	18.2	18.3	17.9	17.7	20.2	19.6	19.6	20.2	21.2	18.7
1959	20.0	23.3	23.9	26.0	24.6	22.6	21.7	23.4	25.5	24.8	28.4	27.6
1960	24.7	28.3	27.4	26.8	27.6	27.6	27.3	27.8	30.5	28.3	31.1	30.1
1961	25.8	27.3	28.8	26.9	27.5	27.7	27.2	25.8	28.4	31.1	32.5	34.0
1962	32.1	33.3	31.3	32.1	30.3	30.6	29.1	26.8	30.8	33.0	34.3	33.2
1963	31.2	32.6	33.4	33.9	32.8	33.8	32.4	29.8	32.9	36.7	37.8	39.7
1964	38.6	39.5	39.7	41.4	39.5	36.9	38.1	34.2	38.5	38.9	44.5	43.9
1965	42.3	40.6	41.7	40.9	40.1	42.7	41.1	34.8	40.9	45.2	48.6	45.7
1966	44.4	47.2	49.3	47.6	47.1	46.8	47.8	45.1	50.1	58.0	61.5	58.0
1967	55.0	59.8	61.2	59.4	55.9	56.3	57.7	49.5	56.5	67.8	67.5	65.1
1968	58.2	72.7	72.4	68.7	67.5	71.1	73.9	63.6	73.2	74.2	77.8	78.5
1969	72.8	77.6	79.4	84.7	78.3	80.3	82.9	70.7	83.5	90.4	88.9	83.5
1970	83.5	85.3	87.1	88.8	83.7	88.5	89.5	78.6	94.2	99.5	99.0	97.1
1971	90.0	95.3	99.5	99.8	94.5	97.7	100.3	87.2	107.0	109.0	114.5	108.4
1972	106.4	114.2	109.5	117.3	114.4	119.2	122.8	100.8	130.6	124.7	129.2	135.1
1973	126.7	140.4	147.4	145.0	140.6	154.9	163.4	122.3	164.8	160.0	164.6	151.6
1974	151.6	157.7	154.8	146.2	160.0	149.6	147.5	122.2	152.6	153.5	147.7	157.9
1975	152.3	173.3	176.1	168.6	150.6	151.3	168.1	120.8	162.7	174.3	179.5	174.8
1976	170.3	180.0	182.3	180.0	168.4	176.4	168.8	141.3	194.1	187.5	186.8	191.4

TABLE A.17

INDEX OF VOLUME OF EXPORTED FOOD (SITC: Section 0)
(1970 = 100)

	JAN	FEB	MARCH	APRIL	MAY	JUNE	JULY	AUG	SEPT	OCT	NOV	DEC
1954	32.2	22.1	27.2	16.3	19.2	16.2	13.3	10.3	49.4	98.2	50.1	39.5
1955	25.1	21.3	17.7	19.7	16.1	16.2	8.7	10.2	57.4	92.9	52.0	32.0
1956	23.7	18.3	16.7	16.3	16.3	15.2	5.5	7.6	36.8	98.0	49.5	48.1
1957	27.9	25.8	29.6	21.1	16.8	15.3	16.2	15.3	76.1	98.8	73.1	64.2
1958	39.2	32.6	25.7	19.5	18.6	17.1	13.3	19.4	71.0	70.0	39.5	52.8
1959	34.3	33.9	23.0	15.1	28.4	41.8	31.4	23.6	65.8	82.0	69.1	62.1
1960	52.6	34.9	36.8	31.0	33.8	30.2	21.1	23.2	68.1	57.5	46.1	54.4
1961	37.8	34.2	31.3	24.8	22.9	23.4	24.1	19.7	82.6	85.7	74.1	81.5
1962	56.7	60.7	45.4	44.0	42.7	37.3	27.4	28.0	92.5	109.2	72.5	87.1
1963	68.4	71.7	59.8	41.6	40.3	38.3	39.9	40.8	47.6	103.1	88.8	57.5
1964	55.5	48.6	45.8	40.1	39.2	30.2	34.9	38.0	57.9	113.0	87.7	107.5
1965	70.1	90.4	81.0	56.3	53.6	37.7	37.9	29.9	60.8	130.8	86.4	102.2
1966	70.6	79.0	64.8	59.2	102.2	73.0	69.8	47.9	96.0	129.1	116.7	172.6
1967	84.0	69.8	53.9	86.6	65.3	62.4	76.8	85.0	170.3	180.5	96.8	105.6
1968	55.9	90.8	71.9	44.8	56.3	46.9	41.6	85.2	137.1	109.9	102.1	126.5
1969	68.2	60.6	52.7	50.0	60.1	69.7	91.1	114.9	86.9	143.4	127.1	135.8
1970	94.8	61.2	71.0	57.0	67.4	106.4	109.1	96.3	109.7	154.3	118.0	153.1
1971	90.2	76.5	60.3	68.5	62.0	75.9	97.8	119.3	152.8	183.9	176.9	176.4
1972	108.7	87.0	115.4	89.9	103.7	133.2	144.7	101.5	144.4	168.8	201.4	246.1
1973	109.0	98.5	111.1	72.8	82.3	106.3	162.7	115.9	122.8	162.0	124.4	167.0
1974	102.9	91.2	83.6	71.9	88.7	141.2	145.0	144.0	106.3	155.4	191.5	219.3
1975	150.8	119.1	99.7	97.4	109.0	205.5	246.5	208.4	163.7	184.7	202.2	341.9
1976	167.2	150.2	123.1	124.5	150.2	202.7	188.0	169.1	200.7	233.5	302.7	390.4

TABLE A.18

INDEX OF VOLUME OF EXPORTED BEVERAGES AND TOBACCO (SITC: Section 1)
(1970 = 100)

	JAN	FEB	MARCH	APRIL	MAY	JUNE	JULY	AUG	SEPT	OCT	NOV	DEC
1954	110.5	68.6	45.4	41.2	38.9	18.9	10.6	8.0	42.2	99.7	201.7	180.2
1955	104.4	66.7	42.8	24.5	38.2	14.8	5.1	8.9	31.8	124.7	224.5	208.4
1956	89.4	42.1	20.2	16.7	14.6	18.6	13.2	10.9	26.0	43.4	166.5	163.5
1957	170.5	105.7	71.6	63.3	42.6	14.8	6.1	5.7	24.0	123.2	229.3	240.3
1958	76.0	49.1	87.6	56.8	48.2	34.7	24.2	16.8	36.9	159.1	279.7	269.8
1959	66.8	68.0	36.5	50.4	34.4	29.8	24.2	6.9	35.5	122.0	181.3	192.6
1960	58.5	75.3	80.0	31.5	37.3	37.3	25.8	22.4	37.6	67.6	235.0	264.1
1961	74.8	100.2	67.7	70.6	43.0	14.2	11.2	12.0	22.1	141.4	263.1	241.6
1962	98.7	125.0	61.3	39.8	22.0	6.2	16.5	15.7	12.9	55.4	153.1	199.6
1963	98.2	108.0	63.8	78.4	60.6	11.4	17.0	9.8	46.1	64.6	227.8	232.0
1964	162.6	103.9	92.3	36.5	35.9	16.2	12.5	7.7	19.0	28.8	362.7	266.8
1965	173.0	110.1	64.6	90.1	85.8	53.8	57.3	19.6	20.2	123.5	222.8	195.8
1966	163.2	151.1	118.8	99.8	76.2	37.8	13.3	11.0	28.8	158.9	196.6	175.8
1967	233.5	154.2	106.5	119.2	61.3	42.5	35.4	17.3	49.6	158.7	245.9	229.4
1968	83.5	71.1	77.8	82.8	61.1	60.1	25.9	29.6	58.5	163.2	176.0	299.1
1969	88.5	97.4	72.8	94.5	133.9	62.0	39.5	26.5	42.7	64.5	161.1	352.6
1970	115.7	73.5	87.3	62.8	87.6	82.9	60.5	58.8	47.3	90.2	208.2	227.6
1971	129.6	90.2	52.0	92.0	52.9	64.0	51.0	53.0	34.4	94.0	220.4	204.0
1972	113.9	199.7	102.0	110.7	83.1	80.1	35.1	52.2	46.6	156.3	187.7	239.7
1973	167.6	98.6	158.1	80.2	50.8	44.6	87.2	76.3	20.8	71.5	52.3	84.6
1974	56.0	73.1	250.6	115.9	221.5	36.9	33.6	79.7	35.4	27.7	141.6	155.3
1975	199.8	209.7	100.9	156.7	45.0	16.0	15.7	7.1	23.7	23.8	146.7	128.4
1976	83.7	96.5	131.6	139.6	130.0	52.9	49.8	40.5	67.3	166.7	197.5	260.8

TABLE A.19

INDEX OF VOLUME OF EXPORTED RAW MATERIALS (SITC: Section 2)
(1970 = 100)

	JAN	FEB	MARCH	APRIL	MAY	JUNE	JULY	AUG	SEPT	OCT	NOV	DEC
1954	28.1	14.8	15.6	15.9	20.4	24.7	18.3	18.0	17.9	13.4	50.8	50.0
1955	34.8	30.7	32.7	39.5	33.7	29.7	28.3	25.2	38.4	79.8	88.2	74.0
1956	70.7	61.0	65.6	61.7	46.4	39.7	35.8	35.0	35.8	83.3	106.3	86.4
1957	68.9	50.7	53.2	48.2	34.5	31.4	28.5	27.2	20.7	36.5	35.1	55.5
1958	75.8	48.1	53.7	59.9	35.8	50.3	62.3	42.7	33.5	51.7	57.7	118.8
1959	94.8	56.9	72.1	67.9	81.8	48.9	39.2	33.6	31.4	71.0	88.4	112.2
1960	71.6	52.0	65.6	57.2	58.2	43.8	32.4	38.5	25.7	40.4	57.0	146.3
1961	38.0	40.8	76.7	77.5	76.7	45.3	54.3	47.9	44.3	46.3	87.7	139.7
1962	124.0	129.0	122.5	92.3	82.4	62.7	39.7	45.3	33.0	52.5	79.7	122.0
1963	100.3	91.2	60.3	91.0	75.3	56.5	74.0	34.0	40.1	48.2	92.2	104.6
1964	110.9	77.7	105.8	117.4	101.1	50.0	56.4	68.3	34.4	32.6	59.7	98.9
1965	66.3	56.2	62.1	77.3	63.0	115.9	56.7	42.8	46.9	47.1	53.8	87.9
1966	89.7	80.6	111.2	88.2	94.1	74.1	53.7	47.7	45.6	59.4	62.5	92.2
1967	128.0	86.0	122.0	75.7	60.1	56.1	51.9	38.0	47.5	58.4	83.3	218.9
1968	90.6	138.0	80.3	135.0	62.6	72.0	52.5	53.2	42.9	43.3	77.8	141.6
1969	68.9	87.1	87.9	91.0	92.8	66.8	71.3	47.0	42.2	76.6	107.1	170.3
1970	93.9	121.2	124.9	101.2	95.3	93.7	81.1	56.2	59.8	78.4	105.0	188.8
1971	103.7	91.4	102.9	147.7	132.8	94.1	91.8	63.4	43.9	55.9	111.3	262.4
1972	110.4	101.9	106.2	104.5	89.7	108.4	79.6	66.5	65.6	64.3	81.3	142.8
1973	84.3	74.3	193.3	120.3	147.1	100.8	111.3	118.9	98.5	91.5	109.0	142.0
1974	102.3	82.2	135.9	102.2	130.7	121.6	66.6	85.4	89.6	88.9	87.3	141.2
1975	112.4	82.0	107.7	95.0	99.5	97.9	97.6	98.7	95.7	70.9	66.5	133.3
1976	106.3	74.2	94.8	119.4	109.7	149.0	76.8	85.8	92.2	112.9	88.5	134.1

TABLE A.20

INDEX OF VOLUME OF EXPORTED CHEMICALS (SITC: Section 5)
(1970 = 100)

	JAN	FEB	MARCH	APRIL	MAY	JUNE	JULY	AUG	SEPT	OCT	NOV	DEC
1954	10.4	14.3	15.5	5.4	9.0	21.8	26.3	20.7	19.3	20.8	17.6	15.3
1955	18.0	11.1	23.2	6.6	15.0	20.2	25.9	33.2	27.1	21.9	21.0	13.6
1956	12.8	18.0	27.8	21.9	6.3	21.5	38.2	39.6	23.5	27.1	20.7	18.5
1957	18.6	17.4	14.9	10.4	7.4	20.1	18.7	19.1	12.2	26.8	22.5	32.1
1958	16.0	25.4	13.7	13.7	8.9	26.2	31.7	18.2	20.2	27.0	18.1	27.4
1959	21.7	20.7	19.6	8.7	6.7	17.7	30.1	29.9	36.3	23.5	15.3	17.1
1960	12.9	16.0	11.2	12.1	3.4	20.2	35.0	22.8	25.8	12.6	13.5	40.2
1961	8.1	13.6	31.2	21.7	13.9	22.0	29.3	23.0	22.4	17.5	33.8	22.5
1962	14.3	25.1	12.9	7.9	9.6	28.4	15.4	24.2	21.0	30.5	11.4	16.4
1963	6.7	15.0	16.6	10.9	6.1	12.5	12.3	12.4	15.0	8.0	16.8	14.0
1964	8.8	12.9	18.2	21.5	8.6	16.4	17.9	13.1	18.2	9.0	32.8	20.4
1965	27.0	30.6	37.0	24.4	11.6	20.4	14.9	12.9	15.9	16.2	12.4	21.4
1966	17.8	16.6	16.9	19.2	14.6	17.6	20.4	7.8	18.6	19.4	31.8	27.7
1967	24.7	22.2	35.1	31.7	28.0	18.8	29.8	79.7	38.8	70.0	87.1	71.8
1968	18.0	106.3	105.6	53.5	121.1	61.6	25.8	98.7	49.8	136.8	76.2	214.0
1969	82.0	44.9	66.6	77.7	63.5	57.0	84.0	76.7	221.7	46.2	84.9	83.7
1970	98.2	62.6	73.4	125.9	146.6	108.0	120.6	48.6	88.9	123.7	74.0	137.2
1971	29.3	111.7	53.2	74.7	88.8	80.2	64.2	225.2	36.0	57.9	164.4	191.2
1972	132.7	118.2	159.0	182.2	50.8	120.5	145.1	111.6	104.8	64.7	208.1	182.4
1973	59.4	132.2	153.6	169.9	105.3	78.9	151.2	182.9	128.0	180.6	143.0	142.6
1974	62.7	129.7	137.9	130.3	132.4	116.7	127.3	152.0	66.2	194.3	82.8	232.8
1975	265.7	109.2	72.5	152.9	190.4	166.1	117.6	115.5	151.0	62.2	113.4	442.9
1976	64.0	86.8	105.4	159.7	140.2	103.9	97.3	109.3	130.9	264.4	144.0	629.4

TABLE A.21

INDEX OF VOLUME OF EXPORTED MANUFACTURES (SITC: Section 6)
(1970 = 100)

	JAN	FEB	MARCH	APRIL	MAY	JUNE	JULY	AUG	SEPT	OCT	NOV	DEC
1954	3.8	2.4	2.9	2.3	2.6	2.0	8.5	1.9	11.1	2.1	2.1	4.1
1955	2.5	5.0	3.0	3.1	3.1	4.2	2.8	3.5	2.3	2.4	2.1	3.3
1956	1.7	4.1	3.2	3.2	2.3	2.7	2.9	3.1	3.7	6.7	6.7	2.0
1957	2.3	3.3	2.3	2.7	3.3	4.3	3.2	3.4	2.4	3.4	3.9	4.4
1958	4.1	4.9	4.7	3.1	3.3	4.2	4.4	3.7	4.4	5.5	2.8	4.6
1959	2.1	2.4	3.0	3.1	4.1	2.4	3.2	3.6	2.8	4.0	3.4	6.1
1960	2.0	3.5	4.5	5.1	3.4	5.1	5.0	3.9	3.5	4.6	5.3	9.6
1961	3.7	2.8	6.3	5.8	5.5	4.0	5.6	9.0	7.1	4.3	6.5	8.6
1962	9.7	5.7	10.7	9.0	6.9	5.8	6.3	9.0	9.2	9.7	8.6	11.7
1963	5.8	8.2	7.9	11.5	8.3	9.0	9.7	6.7	11.6	8.2	13.3	10.4
1964	8.2	12.6	9.7	14.1	9.2	7.5	9.9	11.0	13.3	12.5	13.4	13.4
1965	8.8	12.5	14.2	10.3	13.5	16.0	14.7	13.0	22.2	20.7	16.4	29.9
1966	26.7	12.7	18.9	16.2	23.8	21.2	34.1	25.8	27.1	23.1	50.1	44.6
1967	51.9	21.6	50.0	41.0	35.4	36.0	56.2	34.0	36.0	35.9	42.5	44.7
1968	29.6	53.7	31.0	48.7	46.5	34.0	56.1	55.8	29.3	43.3	41.8	46.2
1969	67.9	72.3	43.5	64.9	53.6	51.9	73.5	87.1	65.2	97.9	75.9	227.9
1970	83.7	87.9	89.5	101.3	95.7	68.8	166.0	77.9	84.4	126.2	107.9	108.8
1971	66.9	58.2	70.5	81.0	85.2	79.5	91.1	108.4	69.9	116.6	121.7	132.1
1972	83.8	94.7	115.9	144.4	158.0	102.2	122.0	156.6	161.7	129.6	159.4	168.3
1973	171.2	175.2	207.3	189.9	169.3	148.4	182.8	240.1	143.7	198.2	306.0	292.2
1974	174.6	180.6	242.1	180.7	305.2	274.5	205.4	213.5	250.3	238.8	291.7	462.5
1975	162.1	154.4	218.5	197.2	184.9	207.7	241.8	204.1	222.6	241.8	208.3	413.7
1976	225.8	224.8	170.2	286.6	224.5	244.3	233.1	262.7	240.1	256.9	254.3	426.4

TABLE A.22

INDEX OF VOLUME OF TOTAL EXPORTS (SITC: Sections 0-9)
(1970 = 100)

	JAN	FEB	MARCH	APRIL	MAY	JUNE	JULY	AUG	SEPT	OCT	NOV	DEC
1954	34.1	23.4	20.4	18.1	17.1	15.1	14.5	10.3	27.8	43.3	55.6	51.7
1955	32.4	25.3	21.0	18.2	19.1	15.2	11.9	12.7	27.3	57.5	66.2	56.4
1956	34.6	25.3	23.0	19.9	15.9	15.9	13.3	13.7	21.4	47.4	61.6	56.9
1957	49.9	34.9	31.1	26.6	19.6	15.5	13.9	13.3	26.1	51.9	63.2	68.8
1958	39.1	29.5	34.5	28.0	21.1	22.6	22.4	18.0	30.5	54.3	66.8	80.5
1959	37.8	31.1	26.4	26.1	28.7	24.8	21.1	16.1	29.2	53.9	62.8	73.6
1960	40.1	34.6	36.4	25.8	26.0	24.5	19.4	18.9	28.5	33.8	61.7	88.5
1961	29.8	33.9	36.8	34.6	28.8	18.6	21.2	19.6	33.3	53.9	81.6	88.1
1962	54.9	59.1	46.4	36.1	32.6	24.9	20.5	23.2	34.7	49.7	63.1	82.2
1963	51.8	55.9	40.3	43.5	35.9	24.6	29.0	20.8	31.4	45.7	80.5	75.4
1964	62.6	47.1	49.8	42.0	35.9	22.2	25.2	27.5	28.3	41.3	99.9	92.9
1965	61.6	55.8	48.1	47.4	44.2	44.2	35.9	25.0	35.0	68.6	74.6	83.8
1966	69.4	66.3	63.6	53.2	61.8	46.7	40.4	32.2	47.3	80.5	91.5	101.7
1967	99.8	65.9	74.1	71.5	54.6	47.0	57.3	48.7	79.3	107.0	103.6	126.9
1968	59.7	83.2	64.1	73.6	63.7	54.9	54.1	63.9	67.9	90.6	90.3	138.7
1969	71.9	75.8	59.5	73.9	78.4	65.4	75.1	75.0	75.2	93.2	109.6	205.0
1970	93.0	81.4	88.2	84.0	92.2	86.5	118.3	73.8	81.4	119.9	126.0	154.7
1971	86.6	78.2	71.0	90.2	84.5	198.6	88.0	104.4	78.1	116.7	157.3	184.2
1972	107.4	118.7	120.9	122.6	117.5	116.4	111.7	109.8	124.1	135.5	171.1	205.0
1973	139.4	129.6	180.6	138.3	129.1	122.2	161.2	172.7	118.2	156.1	175.2	339.2
1974	125.9	131.7	192.9	154.9	212.1	198.6	145.8	165.4	152.7	176.0	215.1	284.2
1975	193.6	162.6	166.0	166.3	144.7	199.8	208.3	184.3	191.2	192.8	209.1	361.9
1976	186.8	191.9	175.0	214.5	216.1	214.3	194.5	189.8	215.8	249.9	256.2	405.5

TABLE A.23

RELATIVE EXPORT PRICES OF FOOD

	JAN	FEB	MARCH	APRIL	MAY	JUNE	JULY	AUG	SEPT	OCT	NOV	DEC
1954	102.9	104.9	103.9	105.0	102.1	103.7	97.9	111.0	112.8	108.1	104.2	118.8
1955	119.8	115.4	122.7	116.0	117.6	117.4	115.9	120.7	119.4	132.6	125.7	122.3
1956	137.7	145.1	142.8	147.8	153.9	153.5	162.0	152.2	143.8	145.8	132.5	129.1
1957	119.5	120.3	123.4	128.3	126.5	127.4	118.2	120.2	124.3	126.1	123.5	120.8
1958	127.5	131.2	131.0	135.7	134.0	129.8	135.2	133.7	144.2	146.3	140.1	133.2
1959	138.4	141.9	137.4	140.9	130.0	141.4	140.4	135.1	131.8	126.4	118.1	105.5
1960	103.0	102.7	110.9	117.0	112.1	112.0	117.3	95.2	104.9	126.3	116.2	103.9
1961	106.8	107.7	108.6	111.3	109.0	105.3	103.2	99.9	103.6	99.4	100.2	98.7
1962	98.8	95.3	96.0	95.4	95.0	93.3	97.3	92.9	101.6	95.4	93.4	95.9
1963	89.8	87.0	91.2	90.3	87.8	95.5	92.2	96.3	109.8	106.6	105.4	102.2
1964	98.4	100.8	102.6	107.9	106.9	112.7	104.7	105.4	111.7	115.0	107.1	99.3
1965	97.0	96.4	97.7	106.1	104.9	109.1	112.1	120.0	105.2	108.8	106.6	98.2
1966	104.8	100.5	168.6	111.3	110.9	107.6	116.5	107.9	112.2	108.4	103.3	95.3
1967	91.1	101.1	107.8	99.6	104.6	104.7	103.1	109.1	110.8	111.9	103.3	103.5
1968	97.0	97.5	100.2	101.8	98.7	103.3	111.4	103.1	113.2	104.6	106.2	103.9
1969	104.2	106.3	105.9	103.4	120.0	102.7	109.4	102.6	105.6	106.6	102.4	95.2
1970	93.7	99.1	96.8	100.1	99.4	106.9	117.0	106.0	93.9	94.7	97.5	96.8
1971	95.6	91.0	96.3	96.6	97.9	108.9	93.6	94.5	105.6	93.7	95.7	95.9
1972	100.2	94.8	92.4	93.5	93.6	96.1	87.7	92.7	95.6	97.8	98.8	99.1
1973	109.6	113.2	122.0	111.4	120.8	141.8	137.8	126.1	130.2	183.7	147.7	125.5
1974	124.4	142.4	137.6	150.1	149.6	150.9	150.8	153.0	163.2	155.8	140.2	119.1
1975	120.8	130.1	135.0	143.2	143.5	150.6	151.9	136.3	131.2	145.9	121.2	127.1
1976	132.4	132.6	130.8	133.9	162.1	147.5	144.7	139.6	141.5	140.2	148.6	140.6

TABLE A.24

RELATIVE EXPORT PRICES OF BEVERAGES AND TOBACCO

	JAN	FEB	MARCH	APRIL	MAY	JUNE	JULY	AUG	SEPT	OCT	NOV	DEC
1954	101.0	115.5	84.4	81.2	92.8	105.0	111.4	117.3	91.7	106.3	121.2	122.6
1955	130.0	126.5	106.6	91.5	91.8	88.4	107.5	109.6	121.8	141.3	142.5	141.1
1956	150.8	174.1	141.3	161.2	145.4	185.0	112.7	145.1	126.1	131.3	166.8	170.7
1957	120.2	118.7	124.8	121.3	116.2	107.7	166.2	104.1	99.4	132.7	123.1	124.7
1958	118.0	129.4	136.8	127.3	115.4	145.2	118.7	109.9	109.2	131.0	128.2	121.3
1959	122.8	118.8	110.9	101.8	98.4	103.3	90.0	145.9	109.9	114.7	128.8	120.4
1960	117.9	124.4	122.9	97.7	102.3	115.7	99.1	99.8	79.3	93.2	113.5	96.9
1961	108.8	110.5	110.7	114.1	82.3	54.0	70.2	82.1	57.4	103.6	124.5	113.0
1962	110.2	135.8	129.8	93.7	105.5	132.3	106.8	111.4	116.3	130.5	126.4	155.3
1963	156.1	156.3	180.6	177.8	179.4	153.8	124.4	108.0	135.7	146.2	143.9	150.0
1964	155.4	151.9	140.4	143.8	146.1	159.1	122.1	114.3	116.4	107.4	127.0	131.1
1965	127.5	120.9	114.2	137.5	136.4	112.6	110.3	129.6	105.8	114.2	109.1	113.7
1966	122.7	131.6	107.7	121.3	102.1	95.0	84.5	101.9	88.3	101.4	106.8	113.3
1967	125.0	124.0	129.7	112.2	105.3	102.1	101.9	98.8	113.2	117.1	116.8	115.9
1968	123.7	115.8	127.0	109.5	100.4	93.7	82.2	101.5	98.6	104.1	98.5	108.7
1969	96.1	97.1	98.0	109.3	135.4	98.6	95.1	84.8	92.7	97.3	105.4	108.1
1970	111.3	92.8	103.7	100.3	94.3	95.6	93.1	92.4	87.7	104.1	106.7	95.2
1971	104.8	107.0	91.1	99.9	80.9	78.3	79.2	88.9	74.1	91.6	97.4	96.7
1972	90.3	106.9	97.5	87.6	87.4	90.6	96.0	80.5	89.4	94.5	89.9	92.3
1973	95.9	100.2	103.8	88.8	95.9	114.3	98.5	109.1	93.6	87.6	92.2	86.6
1974	100.0	109.6	128.3	129.4	124.3	113.3	101.4	97.4	108.3	114.9	103.6	133.2
1975	124.8	119.8	119.0	133.7	121.1	135.1	100.8	260.2	138.4	122.2	157.2	130.5
1976	146.8	109.9	129.7	128.7	132.9	114.3	84.7	110.2	136.7	98.4	127.6	105.5

TABLE A.25

RELATIVE EXPORT PRICES OF RAW MATERIALS

	JAN	FEB	MARCH	APRIL	MAY	JUNE	JULY	AUG	SEPT	OCT	NOV	DEC
1954	113.1	131.9	124.2	114.2	103.4	118.7	115.1	109.9	115.7	114.1	120.5	120.0
1955	115.5	118.8	118.3	104.3	108.9	117.3	116.3	119.3	109.9	103.4	104.1	100.1
1956	107.1	104.3	97.7	104.3	104.9	112.6	117.0	124.4	118.9	99.8	95.5	101.7
1957	103.2	114.7	106.4	109.1	116.4	115.4	125.3	121.0	122.8	113.4	111.0	104.6
1958	100.9	106.7	108.4	104.1	128.9	112.6	106.8	126.3	114.4	101.9	89.2	87.4
1959	81.6	88.5	96.2	86.6	86.7	99.5	123.6	116.1	108.3	92.5	87.9	84.1
1960	82.9	84.7	97.5	103.7	108.8	111.1	112.6	114.7	119.2	107.9	90.1	85.8
1961	97.3	96.7	99.5	100.2	101.3	100.4	101.1	99.5	96.6	96.9	98.8	97.5
1962	97.3	98.4	99.0	99.7	103.3	106.1	102.4	102.4	102.5	101.5	94.8	96.1
1963	94.3	96.6	97.9	97.7	99.1	100.3	101.3	104.0	105.0	98.1	93.9	94.2
1964	93.7	96.5	92.0	97.4	91.8	103.4	107.5	106.3	102.9	101.0	101.1	95.0
1965	100.7	99.5	97.4	101.5	108.3	98.3	106.1	106.4	114.1	99.4	109.4	96.6
1966	94.5	94.9	96.9	99.6	95.0	109.5	111.6	114.2	121.0	113.3	100.5	94.1
1967	93.3	92.4	98.7	107.0	109.5	116.6	119.0	115.7	117.2	104.1	97.0	95.3
1968	96.4	96.2	105.1	106.7	112.8	111.4	126.7	123.2	117.0	117.8	103.5	100.6
1969	98.7	102.6	101.4	103.5	117.1	123.6	132.5	127.0	127.3	114.9	97.4	86.9
1970	100.4	88.7	93.8	97.9	103.4	99.4	118.1	112.9	116.8	107.7	96.2	92.1
1971	93.7	91.6	107.2	91.5	100.6	113.7	104.4	124.0	124.8	110.5	99.8	100.1
1972	107.6	118.3	122.2	111.4	117.5	113.0	108.5	115.4	111.7	111.3	126.9	113.3
1973	108.0	111.3	119.2	123.7	122.6	121.3	126.7	120.2	141.1	124.9	141.3	144.8
1974	133.6	128.6	131.0	100.4	120.7	112.7	108.5	95.6	101.3	100.0	105.8	106.2
1975	98.8	108.5	112.6	114.8	116.6	124.3	114.9	116.6	128.0	135.0	120.0	129.6
1976	122.7	124.5	124.8	123.3	136.7	123.3	117.4	120.6	125.9	120.6	130.5	134.9

TABLE A.26

RELATIVE EXPORT PRICES OF CHEMICALS

	JAN	FEB	MARCH	APRIL	MAY	JUNE	JULY	AUG	SEPT	OCT	NOV	DEC
1954	57.5	82.9	85.3	85.9	78.2	85.7	87.8	87.5	86.0	80.1	77.2	79.2
1955	78.3	81.3	76.4	78.1	71.7	71.4	70.2	69.7	76.1	72.9	71.9	76.2
1956	94.6	83.8	84.6	83.9	78.2	81.8	82.0	80.9	78.5	78.3	78.1	77.4
1957	76.5	74.8	75.6	79.0	74.0	75.5	71.2	71.7	69.8	66.4	69.6	68.5
1958	75.8	72.4	75.2	75.8	70.2	66.3	67.2	65.0	67.3	66.0	63.0	78.5
1959	66.4	64.1	71.7	68.6	68.7	66.7	68.8	73.2	72.0	80.3	82.0	89.2
1960	90.3	93.5	93.4	97.7	82.7	95.3	106.0	114.3	128.0	106.6	123.2	109.4
1961	112.3	119.2	109.5	116.3	103.5	103.5	94.3	100.2	89.5	85.9	84.1	92.6
1962	92.0	86.0	97.3	101.5	93.5	81.9	75.9	83.5	68.2	73.6	95.6	75.4
1963	76.0	69.8	69.3	100.5	89.6	78.1	74.4	75.3	69.2	79.6	71.4	85.9
1964	82.4	78.0	78.5	81.0	91.4	82.2	82.3	123.3	77.7	87.7	72.1	90.5
1965	71.0	73.5	73.5	82.8	97.9	81.3	80.5	90.4	75.2	80.9	77.2	81.6
1966	87.2	78.9	84.5	93.7	98.2	80.5	81.6	136.1	86.6	90.5	78.7	84.9
1967	83.0	77.9	85.2	98.9	107.7	81.7	81.3	67.9	95.5	97.2	59.9	67.5
1968	102.5	93.0	78.0	104.8	80.7	91.0	77.7	89.2	76.0	72.1	72.2	68.0
1969	53.0	63.7	75.4	63.4	93.2	81.3	96.0	91.6	91.9	92.6	98.5	100.8
1970	102.6	92.3	103.8	94.8	95.1	82.8	100.0	107.3	114.6	98.8	107.6	94.0
1971	124.8	112.0	97.7	96.5	81.2	106.3	111.9	109.2	111.0	118.7	101.8	82.7
1972	92.9	102.7	99.3	96.8	126.5	101.3	91.2	102.4	104.2	111.3	99.4	95.0
1973	128.3	111.5	105.9	104.3	119.0	105.2	118.1	115.1	109.6	90.5	110.2	112.4
1974	104.5	105.0	97.8	101.1	100.8	115.4	85.0	102.4	145.7	106.5	143.8	112.7
1975	106.0	115.0	111.1	112.3	115.2	95.8	107.8	118.7	118.1	113.6	117.9	120.9
1976	98.8	115.7	102.6	124.6	115.1	118.3	112.5	111.3	110.8	110.3	119.6	109.6

TABLE A.27

RELATIVE EXPORT PRICES OF MANUFACTURES

	JAN	FEB	MARCH	APRIL	MAY	JUNE	JULY	AUG	SEPT	OCT	NOV	DEC
1954	157.7	148.2	149.4	143.3	147.1	157.0	180.1	166.8	131.9	140.8	157.4	149.0
1955	138.5	146.7	138.5	171.6	147.5	154.0	142.3	146.5	133.2	147.1	146.6	143.6
1956	190.4	163.0	156.4	175.0	162.5	170.5	186.0	157.5	140.4	165.7	146.0	146.8
1957	129.6	139.6	139.6	145.2	149.8	143.3	147.9	161.6	124.4	162.5	166.2	140.5
1958	148.9	150.2	130.0	123.2	124.2	126.1	138.7	105.8	120.4	113.5	113.2	122.4
1959	103.0	127.7	102.0	116.6	123.2	116.9	117.5	116.3	137.0	130.9	125.0	115.9
1960	128.2	118.2	129.1	119.2	133.7	117.0	102.2	121.1	164.6	118.6	120.5	107.5
1961	101.8	112.0	116.3	105.6	100.8	106.7	111.3	122.9	106.6	120.5	118.7	112.5
1962	114.0	102.4	109.7	110.7	108.0	119.1	110.6	112.4	111.6	99.5	114.7	96.1
1963	98.2	111.3	116.2	132.3	119.5	121.5	111.2	107.3	125.1	124.6	114.2	102.5
1964	106.2	110.1	134.2	104.7	107.3	114.6	97.6	107.5	106.7	106.3	101.0	108.2
1965	102.3	114.3	108.8	114.3	100.4	111.4	97.0	105.0	105.2	102.4	107.1	92.1
1966	102.4	124.4	119.3	101.7	107.4	107.9	112.2	118.3	126.9	125.7	107.7	101.1
1967	99.1	120.3	105.8	112.1	103.8	99.5	99.1	103.0	106.1	108.5	103.6	105.5
1968	98.1	105.3	114.1	102.1	102.8	106.7	104.0	101.1	115.9	111.4	107.0	113.0
1969	111.1	113.8	110.3	111.6	112.0	108.5	97.8	101.3	102.8	100.6	99.4	87.3
1970	101.9	108.2	102.8	101.6	96.9	98.4	97.8	100.7	104.2	102.2	95.9	93.0
1971	94.4	99.2	94.5	90.7	92.9	93.1	87.4	91.7	92.7	92.5	88.5	89.1
1972	89.2	87.2	92.5	83.9	83.6	94.0	85.7	78.2	81.2	87.7	86.9	87.1
1973	79.0	87.9	88.7	93.3	83.5	85.5	89.7	75.1	98.8	91.4	80.3	88.1
1974	100.0	111.4	110.4	112.8	99.0	103.4	118.9	118.5	107.8	119.7	101.4	91.2
1975	116.7	117.1	112.0	112.7	112.3	111.4	100.0	104.9	112.1	123.5	115.5	119.1
1976	120.3	125.9	119.1	120.8	121.6	129.0	128.3	128.2	125.1	134.5	124.6	128.7

TABLE A.28

RELATIVE PRICES OF TOTAL EXPORTS

	JAN	FEB	MARCH	APRIL	MAY	JUNE	JULY	AUG	SEPT	OCT	NOV	DEC
1954	105.6	116.0	102.5	98.7	100.9	111.4	115.5	113.2	107.4	110.7	119.1	121.4
1955	122.3	120.1	112.2	105.3	104.5	110.1	111.2	113.3	116.9	129.3	130.2	128.8
1956	131.6	133.6	116.9	130.1	122.7	135.8	119.8	126.1	128.5	128.1	135.9	142.5
1957	117.8	120.0	120.2	119.9	120.5	118.9	124.1	118.3	121.1	128.7	124.8	122.1
1958	114.3	120.5	125.0	117.5	122.6	121.7	110.1	118.7	129.4	130.5	124.8	114.0
1959	104.5	112.0	105.2	97.3	101.5	114.1	116.7	118.5	119.5	113.3	115.5	104.7
1960	100.4	103.8	111.1	105.7	108.1	112.1	110.9	107.1	107.7	111.5	111.2	95.4
1961	104.8	106.4	107.2	108.1	99.1	98.4	101.4	102.8	97.5	100.3	111.9	105.0
1962	103.0	108.9	107.3	99.6	103.0	103.9	101.2	100.9	102.9	103.0	108.1	120.4
1963	115.3	116.3	121.1	125.7	123.4	106.9	103.3	101.9	115.8	118.5	123.7	126.6
1964	125.0	120.9	114.0	109.3	107.0	116.2	107.8	110.2	110.3	113.3	117.4	115.2
1965	113.1	107.4	103.4	117.6	119.0	107.3	109.2	112.5	109.3	111.6	110.3	106.2
1966	112.0	117.4	108.2	114.0	108.1	109.0	113.6	115.1	116.5	110.6	106.2	103.8
1967	108.1	111.8	111.0	108.8	107.6	106.7	104.4	104.5	113.8	113.7	106.1	105.4
1968	106.7	103.4	108.9	107.2	100.9	104.3	107.2	104.7	112.8	104.1	102.9	103.9
1969	99.7	104.0	102.5	104.1	112.7	107.0	108.5	103.8	103.1	105.6	103.4	96.0
1970	102.8	98.3	99.9	99.1	97.8	100.2	103.8	103.6	101.3	99.9	99.6	94.2
1971	98.3	98.8	98.0	93.8	93.0	99.7	92.9	98.0	102.3	96.0	95.7	95.1
1972	95.0	100.3	98.1	91.6	93.4	97.9	90.9	89.4	92.3	97.1	96.8	97.2
1973	96.7	100.3	105.2	102.9	105.5	109.4	111.0	105.4	116.7	120.9	105.5	136.3
1974	124.9	123.8	126.6	141.1	117.9	128.4	125.0	124.2	114.7	121.6	122.4	107.8
1975	128.6	116.2	118.0	123.0	119.7	134.0	125.4	129.0	128.9	123.6	128.1	128.9
1976	123.9	126.0	127.8	127.3	125.1	128.5	131.5	127.3	131.2	129.3	129.8	125.0

TABLE A.29

INDEX OF INDUSTRIAL PRODUCTION OF O.E.C.D.
(1970 = 100)

	JAN	FEB	MARCH	APRIL	MAY	JUNE	JULY	AUG	SEPT	OCT	NOV	DEC
1954	40.0	40.7	42.0	42.0	43.3	42.7	39.3	38.7	44.7	46.0	46.7	46.0
1955	44.0	44.7	46.0	46.0	46.7	47.3	44.0	44.7	48.0	49.3	50.0	48.7
1956	48.0	48.7	49.3	49.3	49.3	50.0	45.3	46.7	50.7	51.3	52.0	50.7
1957	50.0	51.3	52.0	50.7	51.3	52.0	48.7	48.7	52.0	52.0	51.3	49.3
1958	48.0	48.7	48.7	48.0	48.0	49.3	46.7	46.7	50.7	51.3	52.7	51.3
1959	50.7	52.7	53.3	54.7	55.3	56.0	51.3	50.7	54.7	56.0	56.0	56.0
1960	55.0	56.0	57.0	57.0	57.0	57.0	53.0	53.0	57.0	58.0	58.0	56.0
1961	54.0	56.0	57.0	58.0	58.0	59.0	56.0	55.0	60.0	61.0	62.0	61.0
1962	59.0	61.0	62.0	62.0	63.0	63.0	59.0	58.0	64.0	64.0	65.0	63.0
1963	62.0	64.0	66.0	66.0	67.0	68.0	63.0	63.0	68.0	70.0	71.0	69.0
1964	68.0	70.0	71.0	72.0	73.0	73.0	67.0	60.0	73.0	74.0	75.0	74.0
1965	73.0	75.0	76.0	76.0	77.0	78.0	72.0	72.0	78.0	81.0	81.0	79.0
1966	78.0	80.0	83.0	82.0	83.0	84.0	77.0	77.0	84.0	86.0	86.0	83.0
1967	81.0	84.0	84.0	85.0	85.0	86.0	79.0	86.0	86.0	89.0	90.0	89.0
1968	85.0	89.0	91.0	89.0	91.0	93.0	86.0	87.0	93.0	96.0	98.0	95.0
1969	93.0	96.0	98.0	98.0	99.0	101.0	93.0	93.0	100.0	102.0	103.0	100.0
1970	96.0	100.0	102.0	102.0	102.0	103.0	95.0	94.0	101.0	103.0	102.0	100.0
1971	99.0	102.0	103.0	103.0	103.0	105.0	97.0	95.0	104.0	107.0	107.0	103.0
1972	102.0	106.0	109.0	110.0	110.0	112.0	102.0	103.0	112.0	117.0	117.0	115.0
1973	112.0	118.0	120.0	120.0	120.0	123.0	113.0	113.0	123.0	126.0	126.0	122.0
1974	117.0	122.0	124.0	123.0	123.0	126.0	114.0	113.0	122.0	122.0	124.0	111.0
1975	105.0	109.0	110.0	109.0	109.0	112.0	103.0	104.0	114.0	117.0	118.0	114.0
1976	111.0	118.0	120.0	121.0	121.0	123.0	113.0	111.0	124.0	125.0	126.0	122.0

TABLE A.30

INDEX OF INDUSTRIAL PRODUCTION OF FOOD,
BEVERAGES AND TOBACCO OF O.E.C.D.
(1970 = 100)

	JAN	FEB	MARCH	APRIL	MAY	JUNE	JULY	AUG	SEPT	OCT	NOV	DEC
1954	47.4	46.6	49.6	51.1	52.6	54.1	50.4	50.4	52.6	60.9	64.7	63.2
1955	49.6	50.4	52.6	51.1	53.4	55.6	54.9	57.1	57.9	59.4	58.6	55.6
1956	52.6	54.1	55.6	54.1	56.4	57.9	57.1	59.4	61.7	63.2	61.7	57.9
1957	54.9	56.4	57.9	56.4	58.6	60.9	60.2	61.7	63.2	63.2	61.7	58.6
1958	56.4	57.9	58.6	57.9	60.2	62.4	61.7	63.9	64.7	66.2	64.7	61.7
1959	58.6	60.2	61.7	60.9	63.2	64.7	63.9	66.9	68.4	68.4	66.9	63.9
1960	62.0	64.0	66.0	64.0	66.0	69.0	67.0	69.0	72.0	74.0	71.0	69.0
1961	64.0	66.0	68.0	66.0	68.0	70.0	68.0	71.0	73.0	75.0	75.0	70.0
1962	67.0	68.0	71.0	69.0	71.0	74.0	72.0	73.0	76.0	77.0	77.0	73.0
1963	71.0	71.0	74.0	74.0	76.0	78.0	76.0	77.0	79.0	82.0	82.0	79.0
1964	75.0	76.0	76.0	78.0	80.0	81.0	79.0	80.0	83.0	86.0	85.0	82.0
1965	77.0	78.0	80.0	80.0	81.0	84.0	81.0	83.0	86.0	88.0	89.0	84.0
1966	79.0	81.0	84.0	83.0	85.0	88.0	85.0	87.0	89.0	92.0	91.0	87.0
1967	82.0	84.0	87.0	88.0	89.0	91.0	89.0	90.0	93.0	95.0	94.0	91.0
1968	85.0	88.0	90.0	90.0	91.0	94.0	91.0	93.0	96.0	99.0	98.0	93.0
1969	90.0	91.0	95.0	93.0	95.0	97.0	94.0	97.0	100.0	102.0	101.0	97.0
1970	93.0	96.0	98.0	98.0	100.0	102.0	98.0	100.0	104.0	106.0	104.0	101.0
1971	98.0	98.0	101.0	101.0	101.0	105.0	102.0	103.0	110.0	109.0	108.0	103.0
1972	100.0	102.0	105.0	105.0	106.0	108.0	104.0	107.0	111.0	114.0	113.0	109.0
1973	102.0	105.0	108.0	109.0	111.0	114.0	109.0	113.0	119.0	120.0	120.0	114.0
1974	108.0	109.0	112.0	113.0	114.0	119.0	111.0	114.0	120.0	120.0	117.0	110.0
1975	103.0	105.0	108.0	110.0	112.0	116.0	112.0	116.0	121.0	123.0	121.0	116.0
1976	109.0	112.0	114.0	117.0	117.0	123.0	118.0	120.0	129.0	129.0	127.0	120.0

TABLE A.31

INDEX OF INDUSTRIAL PRODUCTION OF CHEMICALS OF O.E.C.D.
(1970 = 100)

	JAN	FEB	MARCH	APRIL	MAY	JUNE	JULY	AUG	SEPT	OCT	NOV	DEC
1954	24.7	24.7	25.8	25.8	26.3	26.3	25.3	25.3	26.9	27.4	27.4	26.9
1955	27.4	28.0	29.0	29.0	29.0	29.6	28.0	28.5	30.1	30.6	30.6	30.1
1956	30.1	30.1	31.2	31.2	31.2	31.2	29.6	29.6	31.7	32.3	32.3	31.7
1957	32.3	32.8	33.3	32.8	33.3	32.8	31.7	31.7	33.9	34.4	34.4	33.3
1958	32.8	33.3	33.3	33.3	33.3	34.4	32.8	32.8	35.5	36.6	36.6	36.0
1959	36.6	37.6	38.2	38.7	39.2	39.8	37.6	37.6	40.3	40.9	40.9	40.3
1960	41.0	42.0	42.0	42.0	42.0	43.0	41.0	41.0	42.0	43.0	43.0	42.0
1961	42.0	43.0	44.0	44.0	45.0	45.0	43.0	44.0	45.0	47.0	47.0	47.0
1962	47.0	48.0	49.0	49.0	50.0	50.0	48.0	48.0	50.0	51.0	51.0	50.0
1963	50.0	52.0	53.0	55.0	55.0	56.0	53.0	54.0	55.0	58.0	58.0	56.0
1964	57.0	58.0	59.0	60.0	61.0	62.0	59.0	59.0	62.0	63.0	63.0	62.0
1965	63.0	64.0	65.0	65.0	66.0	67.0	64.0	65.0	67.0	69.0	68.0	67.0
1966	68.0	69.0	72.0	72.0	73.0	74.0	69.0	70.0	73.0	75.0	75.0	74.0
1967	74.0	74.0	76.0	75.0	76.0	77.0	73.0	74.0	78.0	81.0	81.0	80.0
1968	80.0	83.0	84.0	84.0	85.0	87.0	83.0	84.0	88.0	90.0	91.0	90.0
1969	89.0	92.0	93.0	94.0	95.0	97.0	92.0	92.0	96.0	98.0	98.0	97.0
1970	96.0	98.0	100.0	100.0	100.0	102.0	99.0	97.0	101.0	104.0	103.0	101.0
1971	100.0	104.0	105.0	106.0	105.0	108.0	103.0	102.0	108.0	111.0	111.0	109.0
1972	109.0	112.0	115.0	115.0	116.0	118.0	112.0	112.0	113.0	123.0	124.0	122.0
1973	121.0	126.0	128.0	128.0	129.0	131.0	125.0	124.0	130.0	134.0	134.0	128.0
1974	128.0	133.0	134.0	135.0	135.0	137.0	129.0	127.0	133.0	132.0	125.0	115.0
1975	112.0	113.0	113.0	114.0	115.0	119.0	114.0	113.0	123.0	128.0	128.0	124.0
1976	123.0	131.0	133.0	135.0	133.0	136.0	130.0	125.0	138.0	138.0	139.0	135.0

TABLE A.32

INDEX OF INDUSTRIAL PRODUCTION OF THE MANUFACTURING INDUSTRIES OF O.E.C.D.
(1970 = 100)

	JAN	FEB	MARCH	APRIL	MAY	JUNE	JULY	AUG	SEPT	OCT	NOV	DEC
1954	38.4	39.1	40.4	40.4	41.7	41.1	38.4	37.1	43.0	43.7	44.4	43.7
1955	42.4	44.4	45.7	45.7	46.4	46.4	43.7	43.7	47.0	49.0	49.0	47.7
1956	47.0	47.7	48.3	48.3	48.3	49.0	45.0	45.7	49.7	51.0	51.0	49.7
1957	49.0	50.3	51.0	50.3	51.0	51.0	47.7	47.7	51.0	51.0	50.3	48.3
1958	47.0	47.7	47.7	47.0	47.7	48.3	45.7	45.7	49.7	51.0	51.7	50.3
1959	49.7	51.7	53.0	53.6	55.0	55.0	50.3	49.7	53.6	55.6	55.0	55.6
1960	54.0	56.0	57.0	56.0	57.0	57.0	53.0	52.0	57.0	58.0	57.0	55.0
1961	53.0	55.0	57.0	57.0	58.0	59.0	55.0	54.0	59.0	61.0	61.0	60.0
1962	58.0	60.0	62.0	62.0	63.0	63.0	58.0	58.0	63.0	64.0	64.0	62.0
1963	61.0	63.0	65.0	66.0	67.0	68.0	62.0	63.0	68.0	70.0	71.0	68.0
1964	67.0	70.0	71.0	72.0	72.0	73.0	67.0	67.0	73.0	75.0	75.0	73.0
1965	72.0	75.0	76.0	77.0	78.0	78.0	72.0	72.0	78.0	81.0	81.0	79.0
1966	77.0	80.0	83.0	83.0	84.0	85.0	78.0	78.0	84.0	87.0	86.0	83.0
1967	81.0	83.0	85.0	85.0	85.0	86.0	79.0	80.0	87.0	89.0	90.0	88.0
1968	85.0	88.0	91.0	90.0	91.0	93.0	86.0	87.0	94.0	97.0	98.0	95.0
1969	92.0	96.0	99.0	99.0	100.0	102.0	93.0	93.0	101.0	103.0	103.0	99.0
1970	96.0	100.0	102.0	102.0	103.0	103.0	95.0	94.0	101.0	103.0	102.0	100.0
1971	98.0	101.0	101.0	101.0	103.0	105.0	97.0	95.0	105.0	108.0	107.0	102.0
1972	101.0	105.0	109.0	110.0	110.0	112.0	102.0	103.0	113.0	118.0	118.0	115.0
1973	111.0	117.0	121.0	120.0	121.0	124.0	114.0	113.0	124.0	127.0	127.0	121.0
1974	117.0	122.0	125.0	124.0	125.0	127.0	115.0	113.0	123.0	123.0	120.0	110.0
1975	104.0	107.0	108.0	109.0	109.0	112.0	103.0	103.0	114.0	118.0	117.0	112.0
1976	109.0	117.0	120.0	121.0	121.0	124.0	114.0	111.0	125.0	126.0	126.0	120.0

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