Trade and Competitiveness between Turkey and the EU:
Time Series Evidence

Utku Utkulu\textsuperscript{\textcopyright} \quad Dilek Seymen\textsuperscript{\textcopyright}

Abstract
The paper basically aims to clarify the level of price competitiveness of the Turkish firms towards the EU Single Market in aggregate level. Thus it naturally examines the demand for exports and imports for Turkey in relation to the EU. In order to model the trade between Turkey and the EU, we employ a time series analysis, namely cointegration method with error correction and causality mechanisms, for the period 1963-2002. The paper also deals with the possible effects of factors such as structural breaks, integration of markets, product innovation, supply, and omitted variables as regards the significance and the magnitude of the income and price elasticities. In the light of our empirical findings, some policy implications are drawn.

Keywords: trade, competitiveness, income and price elasticities, EU, Turkey, time series modeling, cointegration.

JEL Classification: F11, F13, F14, F15

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1. Introduction

The aim of this paper is to analyse the behaviour of the Turkish and the EU trade relationship by modelling exports and imports to understand the nature and the driving forces of the competition. In doing this, we focus on the estimation of aggregate export and import demand functions and hence on the determination of the relevant price and income elasticities, among others. It is well-known that the effectiveness of foreign trade policy is dependent on the significance and the size of the income and price elasticities of exports and imports (Goldstein and Khan, 1985). Traditionally, most researchers have assumed that foreign trade is determined by either income or price effects (or combination of both). This orthodox view has been influenced by traditional theories of balance of payments such as the elasticity and absorption approaches. These theoretical models emphasize the role of ‘price’ and ‘income’ effects in foreign trade. The essence of the elasticity approach is embodied in the famous Marshall-Lerner condition.\(^1\)

Much of the recent debate focuses on the following question. What is the driving force behind the rapid growth of the NIC exports? The answer of this question lies in the estimated income and price elasticities which emerged from empirical studies of the demand for exports. In other words, the key point of the debate is the issue of whether this rapid export growth of some LDCs is to be regarded as reflecting high and statistically significant price elasticities, high and statistically significant income elasticities, or both.\(^2\)

Trade modeling is an effective method to understand the driving forces behind the competition. However, non-price factors are not taken into account by the orthodox trade models. Here in this paper we model the trade between Turkey and the EU by not only taking classical price and income effects but also including non-price factors such as product innovation, commodity composition effects, integration of markets and etc. This paper empirically models and investigates the EU-Turkey trade by using time series techniques. The organisation of the paper is as follows. Section 2 presents and discusses the review of the empirical literature as far as income and price effects of Turkey are concerned. Section 3 provides an evaluation of the Turkey-EU trade relations and post-liberalisation (i.e. post-1980) period of the Turkish foreign trade. A trade model is presented in Section 4. Section 5 describes the econometric methodology employed. Section 6 provides the data and reports the empirical results of the application of the theoretical model outlined to the EU-Turkey trade. The last section offers some conclusions and implications.

2. Some Recent Estimates of Income and Price Elasticities for Turkey

Utkulu (1995), using cointegration analysis, estimated long-run price estimates for Turkish exports and imports as \(-3.12\) and \(-0.41\) respectively. Low but significant income effects are also reported suggesting that both income and price effects have been the driving forces. Kotan and Saygılı (1999) estimate import demand function for Turkey by employing Engle-Granger method. Using quarterly data for the period 1987-1999, they found that import demand is price and income inelastic in the long run. The income elasticity of import demand was found 0.26, whereas exchange rate elasticity and domestic price elasticity of import was found 0.24 and 0.37 respectively. Sahinbeoğlu and Ulaşan (1999), estimated export supply and demand function for Turkey using 1987-1998 sesional data. According to their model estimations, both price and

\(^1\) The simple Marshall-Lerner condition states that, given a balanced current account, a devaluation will improve the balance of payments on current account, if and only if the sum of the price elasticities of domestic demand for imports plus foreign demand for exports exceeds unity (in absolute terms). Given the rather strong assumptions of the simple Marshall-Lerner condition, any result should be treated as ‘indicative’ rather than ‘conclusive’ unless a modified version of the simple Marshall-Lerner is implemented.

\(^2\) The conventional explanation is that although price elasticities of demand for NIC exports are low (or insignificant), the world income elasticity of demand for the NICs’ exports appear to be significant and high. One possible answer to the question of how it is that the NICs have found themselves in a position of continuously facing highly income elastic demand curves comes from Krugman (1989) who suggests that the growth process of the NICs has been driven by a continuous process of product innovation and diversification. Krugman recognizes that the close relation between growth rates and the relative size of income elasticities\(^2\) could have two types of interpretation. On the one hand, income elasticities could determine growth by imposing a balance of payments constraint on demand. Differential growth rates could affect trade flows in such a way as to create obvious differences in income elasticities, on the other. Krugman dismisses the first explanation that growth may be demand-constrained by the balance of payments. Instead, he argues that faster growth in one country leads to a greater supply of exports. Accordingly, as a country’s relative growth rate changes, its apparent income elasticities change too, maintaining the 45-degree rule.
income elasticities of real export demand (0.43 and 0.10 respectively) and supply function estimates are also inelastic. The estimation results indicate that in analyzing exports for the period after 1994, traditional export equations are not sufficient for forecasting and policy simulations. Variables such as uncertainty indicator, or investment have crucial roles in explaining exports. Senhadji and Montenegro (1999) estimated export demand elasticities for a large number of developing and industrial countries including Turkey. According to their results, for Turkey the long run price elasticities of demand is far greater than one (-4.72), i.e. very price elastic, whereas foreign income elasticity of demand less than one (0.51). Cosar (2002) estimated the export demand elasticities of foreign income, real exchange rate and sectoral production by using panel unit root and cointegration test for the period of 1989I-2000IV. According to the estimation results, the real exchange rate elasticity of total export demand is found to be less than one (0.42), whereas the income elasticity is found to be greater than one(4.5). The production elasticities of sectoral export demand are found greater than one for most of the manufacturing goods accept machinery.

Neyapti et al. (2003) estimated export and import functions of Turkey with the EU and non-EU Countries using panel data set. First they estimated income and price elasticities of exports and imports considering Turkish bilateral export. They measure income (2.0) and price elasticity (-0.66) (with respect to real exchange) of exports are both significant. In accordance to their estimation, Turkish bilateral exports are income elastic but price inelastic. Similar results are obtained in the estimation of import function. While domestic income elasticity is estimated more than unity, relative price elasticity is found less than one. Then adding the customs union (CU) effect to the model, they observed that, while income elasticity of both exports and imports are lower for the EU countries (and especially in the CU period), the effect of real exchange rate on Turkey’s exports to the EU is stronger for the CU period (though not earlier). For imports they estimated that real appreciation of TL has had a positive impact on imports especially from the EU countries, though not in the CU period. According to the study custom union agreement and resulting change in the tariff structure caused a significant change in the direction of trade towards the EU and away from the non-EU countries.

3. Turkey and the EU Trade Relationship

A Brief Look at the Turkish Experience

A turning point in Turkish trade policy came in January 1980. Inward-looking (import substitute) industrialization strategy was replaced by an outward-oriented (export-led) growth strategy which relied on more market-based economy. Gradual import liberalization, more flexible exchange rate regime, more effective export promotion to encourage rapid export growth were general objectives of this reform program. There is little doubt that the Turkish economy has achieved an impressive transformation from an inward-looking economy to an outward-oriented one (see Table 1).

Table 1: Some Key Trade Indicators of Turkey (1963-2002) (Billion US dollars)

<table>
<thead>
<tr>
<th>Year</th>
<th>Exports (Trillion)</th>
<th>Imports (Trillion)</th>
<th>Trade Deficit (Trillion)</th>
<th>Exports/Import (%)</th>
<th>Export/GDP (%)</th>
<th>Imports/GDP (%)</th>
<th>TRGDP/Total Exp (%)</th>
<th>Manu.Exp/Total Export (%)</th>
</tr>
</thead>
<tbody>
<tr>
<td>1963</td>
<td>368</td>
<td>687</td>
<td>-319</td>
<td>53.5</td>
<td>4.9</td>
<td>9.2</td>
<td>0.248</td>
<td>19.8</td>
</tr>
<tr>
<td>1980</td>
<td>2.910</td>
<td>7.909</td>
<td>-4.999</td>
<td>36.8</td>
<td>4.1</td>
<td>11.2</td>
<td>0.152</td>
<td>36.0</td>
</tr>
<tr>
<td>1985</td>
<td>9.088</td>
<td>11.343</td>
<td>-3.435</td>
<td>50.2</td>
<td>11.8</td>
<td>16.9</td>
<td>0.437</td>
<td>75.3</td>
</tr>
<tr>
<td>1990</td>
<td>12.959</td>
<td>22.302</td>
<td>-9.343</td>
<td>60.6</td>
<td>18.6</td>
<td>14.8</td>
<td>0.437</td>
<td>79.0</td>
</tr>
<tr>
<td>1995</td>
<td>21.223</td>
<td>35.708</td>
<td>-14.071</td>
<td>55.2</td>
<td>21.2</td>
<td>14.8</td>
<td>0.440</td>
<td>88.2</td>
</tr>
<tr>
<td>1996</td>
<td>26.261</td>
<td>43.627</td>
<td>-17.366</td>
<td>54.1</td>
<td>21.2</td>
<td>14.8</td>
<td>0.447</td>
<td>87.1</td>
</tr>
<tr>
<td>2000</td>
<td>26.974</td>
<td>48.559</td>
<td>-21.585</td>
<td>58.7</td>
<td>21.2</td>
<td>14.8</td>
<td>0.447</td>
<td>88.5</td>
</tr>
<tr>
<td>2001</td>
<td>31.334</td>
<td>54.503</td>
<td>-22.169</td>
<td>65.3</td>
<td>21.2</td>
<td>14.8</td>
<td>0.438</td>
<td>91.2</td>
</tr>
<tr>
<td>2002</td>
<td>35.762</td>
<td>51.270</td>
<td>-15.508</td>
<td>75.7</td>
<td>21.2</td>
<td>14.8</td>
<td>0.512</td>
<td>93.1</td>
</tr>
</tbody>
</table>

Source: State Planning Organisation (SPO), IMF Financial Statistics, several years.

Turkey's export performance has been impressive, especially in the first half of the 1980s. 1987-89 period had witnessed relatively small increases in export. Between 1980 and 1990, exports grew at an average annual rate of 17.2%, while manufactured goods exports increased in current US dollars at an
annual rate of 26.2%. Exports came from 2.9 billion US dollars in 1980 to 12.9 billion US dollars in 1990, and the export/GDP ratio increased. The export composition changed in favour of manufactured goods. In addition to the leading subsectors like textiles and clothing, iron and steel, several other subsectors also enjoyed remarkable expansion. Along with the manufactured sectors, many service export industries such as tourism, transportation and contracting also expanded their shares.

Turkey’s export performance slowed significantly especially during the 1989-1993 period due to the expansionary monetary policies and the appreciation of the Turkish lira. A stabilization program was announced in 1994 with the aim to reduce the domestic demand and rate of inflation and to increase exports through the real depreciation of the Turkish lira. As a result of the program, exports increased in this period. The growth tendency of exports continued till 1997 when the export performance decreased due to the crisis in the Southeast Asia and the Russian Federation. The earthquakes occurred in 1999 also affected the economic conditions negatively (Coşar, 2002). In addition, Turkey liberalised its import regime substantially from 1980 onwards. In short, nominal tariff rates were reduced remarkably, quantitative restrictions were abolished, and bureaucratic controls over imports were also relaxed continuously.

The Turkish economy has witnessed a new recession recently (in 2001). The major challenge facing the new government is to put the macroeconomic balances in order, to be able get rid of the ongoing recession, also to establish a credible strategy for achieving sustainable internal and external deficits, lower inflation and sustainable economic growth in the medium term. (Utkulu and Özdemir, 2003).

The **EU and Turkey trade relationship**

There exists two basic dimensions of the European Union (EU) and Turkey relationship (Seymen, 1998a). The first began with Turkey’s application as an associate member to the European Economic Community (EEC) in 1959. This application forms the basis of Turkey’s current Customs Union (CU) Relations. The second is the application for full membership to the EC in 1987. This study focuses particularly on the association relationship between parties to see trade relations in specific.

Economic relations between two parties have been strong since the early 1950s, but were intensified over recent decades. The long-standing preferences between Turkey and the EU have resulted in the EU being not only the most important market for Turkey (50.5% of Turkey’s exports in 2002) but also one of the main sources for imported goods (45.1% of Turkey’s imports in 2002). The Community accounts for nearly half of Turkey’s total imports and exports as compared to other partners (See Table 2).

**Table 2: Turkey and the EU Trade**

<table>
<thead>
<tr>
<th>Year</th>
<th>Export (TR)</th>
<th>Change (%)</th>
<th>Export (to the EU)</th>
<th>Change (%)</th>
<th>EU Share of Export</th>
</tr>
</thead>
<tbody>
<tr>
<td>1995</td>
<td>21.6</td>
<td>11.1</td>
<td>51.2</td>
<td></td>
<td></td>
</tr>
<tr>
<td>1996</td>
<td>23.2</td>
<td>7.3</td>
<td>49.7</td>
<td></td>
<td></td>
</tr>
<tr>
<td>1997</td>
<td>26.3</td>
<td>13.1</td>
<td>46.6</td>
<td></td>
<td></td>
</tr>
<tr>
<td>1998</td>
<td>27.0</td>
<td>2.7</td>
<td>50.0</td>
<td></td>
<td></td>
</tr>
<tr>
<td>1999</td>
<td>26.6</td>
<td>-1.4</td>
<td>54.0</td>
<td></td>
<td></td>
</tr>
<tr>
<td>2000</td>
<td>27.8</td>
<td>4.5</td>
<td>52.2</td>
<td></td>
<td></td>
</tr>
<tr>
<td>2001</td>
<td>31.3</td>
<td>12.8</td>
<td>51.4</td>
<td></td>
<td></td>
</tr>
<tr>
<td>2002</td>
<td>35.8</td>
<td>14.1</td>
<td>50.5</td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

<table>
<thead>
<tr>
<th>Year</th>
<th>Import (TR)</th>
<th>Change (%)</th>
<th>Import (from the EU)</th>
<th>Change (%)</th>
<th>Import share from the EU</th>
</tr>
</thead>
<tbody>
<tr>
<td>1995</td>
<td>35.7</td>
<td>16.9</td>
<td>47.2</td>
<td></td>
<td>-5.8</td>
</tr>
<tr>
<td>1996</td>
<td>43.6</td>
<td>22.2</td>
<td>37.2</td>
<td>53.0</td>
<td>-11.6</td>
</tr>
<tr>
<td>1997</td>
<td>48.6</td>
<td>11.3</td>
<td>24.9</td>
<td>7.5</td>
<td>51.2</td>
</tr>
<tr>
<td>1998</td>
<td>45.9</td>
<td>-5.4</td>
<td>24.1</td>
<td>-3.2</td>
<td>52.4</td>
</tr>
<tr>
<td>1999</td>
<td>40.7</td>
<td>-11.4</td>
<td>21.4</td>
<td>-11.0</td>
<td>52.6</td>
</tr>
<tr>
<td>2000</td>
<td>54.5</td>
<td>34.0</td>
<td>26.6</td>
<td>24.3</td>
<td>48.8</td>
</tr>
<tr>
<td>2001</td>
<td>41.4</td>
<td>-24.0</td>
<td>18.3</td>
<td>-31.3</td>
<td>44.2</td>
</tr>
<tr>
<td>2002</td>
<td>51.3</td>
<td>23.8</td>
<td>26.5</td>
<td></td>
<td>45.1</td>
</tr>
</tbody>
</table>

**Source:** State Planning Organisation (SPO), IMF Financial Statistics, several years.

The scope of the CU excludes Turkey from some of the crucial aspects of the common market: the common agricultural policy, including the free circulation of agricultural products; the free movement of labour and capital; and moves towards a single currency. Unlike countries in the European Economic Area,
Turkey may also be subject to anti-dumping measures by the EU. The financial support originally envisaged from the EU to Turkey has not yet been made available (Hartler and Laird, 1999).

The CU Decision caused some changes in Turkish trade (Seymen, 1998b). Turkey's imports from the EU in 1996 (the first year of the CU implementation) reached $23 billion, with an increase of 37.2%. Considering the 22.2% increase in Turkish total imports in 1996, it is clear that the CU had a certain impact on the increase in imports. Turkey's export to the EU totalled $11.5 billion with an increase of 4.2%, below the 7.3% increase in total exports in 1996. Consequently, Turkey's foreign trade deficit with the Union doubled and increased to $11.6 billion in 1996. In 1999 and 2001 economic stagnation affected Turkey's trade negatively, so imports from the EU decreased as well. With the exception of periods of economic crises, increases in imports was greater than the exports increases. So resulting trade deficit was high. In 2002 Turkey's exports to the EU is $18.1 billions with the %12 increase. Turkey's import from the EU is $23.1 billions with the %26.5 increase. Turkey's trade deficit with the Union is $5.1 billions in 2002.

Figures in Table 2 suggest that the EU share in the Turkish exports and imports have always been around 50 per cent. This shows that Turkey and the EU have been traditional and stable trade partners over time. This fact has not changed even in the years of economic crises of 1999, 2000 and after. Empirically, it is difficult to measure the effects of the CU (trade creation and trade diversion effects-revenue lose-sectoral effects) in such a short period of time. Instead, in our study, we employed trade modeling for the period 1963-2002, which produced relevant price and income elasticities that are crucial for policy implications for both Turkey and the EU.

4. A Simple Trade Model

We consider an imperfect substitutes model of trade, the key underlying assumption of which is that neither exports nor imports are perfect substitutes for domestic goods (see Goldstein and Khan, 1985). By definition, we have

\[ TBS_t = X_S_t - M_S_t \]  \hspace{1cm} (1)

where TBS, X_S and M_S are total merchandise trade balance, exports and imports in US dollars respectively. Let us now define export and import volumes as

\[ X_V_t = X_S_t/P_X_t \]  \hspace{1cm} (2)
\[ M_V_t = M_S_t/P_M_t \]  \hspace{1cm} (3)

where X_V, M_V, P_X and P_M are export volume, import volume, export price index and import price index in US dollar terms.

It is now our aim to develop export and import demand models, and to estimate the corresponding income and price elasticities for foreign trade of Turkey. The conventional long-run export and import demand functions are as follows:³

³As pointed out by Goldstein and Khan (1985), when the two-country model is left for the n-country real world, the symmetry between export and import demand equations disappears. This is due to the fact that a country's total imports face competition only from domestic producers, while a country's total exports face competition not only from domestic producers in the export market but also from 'third country' exporters to that market. Indeed, the traditional practice in specifying export demand equations is to assume that the major price competition occurs among exporters. Evidence by Arslan and van Wijnbergen (1993) suggests that the Turkish export firms compete with the firms from third countries exporting to the same market. We specify the variables in logarithms so that the coefficients are the relevant relative price and income elasticities. In what follows, all variables are also in logarithms unless otherwise stated. For the choice of the functional form (i.e. linear versus log-linear), see, e.g., Khan and Ross (1977). The evidence suggests that log-linear specification is preferable (see, e.g., Goldstein and Khan, 1985). We also assume that both export and import demands are homogenous of degree zero in prices (i.e. relative price restriction). For information about homogeneity postulate, see, e.g., Leamer and Stern (1970), Murray and Ginman (1976) and Goldstein et al. (1980), among others.
EXPORT DEMAND: \( XV = f_1 \left[ \left( \frac{PX}{PW} \right), \ YW \right] \)  
(-) (+)  

IMPORT DEMAND: \( MV = f_2 \left[ \left( \frac{PM}{PD} \right), \ YD \right] \)  
(-) (+)  

where \( XV \) and \( MV \) represent the volumes of export and import goods respectively; \( \left( \frac{PX}{PW} \right) \) represents the relative export prices, i.e., the ratio of export prices of the exporting country to world prices expressed in common currency units which we call export price competitiveness; \( \left( \frac{PM}{PD} \right) \) is the relative import prices, i.e., the ratio of import prices facing the importing country to domestic prices (preferably wholesale price index) expressed in common currency units which we call import price competitiveness; \( YW \) is a scale variable which captures world demand conditions; and \( YD \) is the real domestic income. The signs given in the parentheses are the expected ones.

Traditional models, see eq. 4 and 5, estimate export and import demand as functions of relative prices and income but omit other factors which might be relevant and statistically significant. In line with the theory, the following can be added to the right hand side of the export demand equation as possible explanatory variables: non-price / supply-side factors such as product types and quality, product and process innovation (Muscatelli et al., 1991; Landesmann and Snell, 1993), and economic integration effects (Madsen and Damania, 1994). In a similar way, the relevant variables which may be included to the import demand equation as explanatory variables are as follows: indicators of import capacity such as external debt stock, foreign exchange inflows, and non-gold international reserves.

In addition, the following theoretical issues have to be addressed as well. First, it is mentioned by Landesmann and Snell (1993) and Muscatelli et al. (1991) that highly aggregated export models such as the one explained above by their nature do not address the effects regarding the transformation on the types and quality of goods produced and exported. That is why, as the composition of a country's aggregate exports change through time, one would expect structural changes in the estimated parameters. Given the relatively small sample at our disposal, rather than assume compositional change effects, we wanted to test directly for their significance using a practical proxy for commodity composition of Turkish exports goods.

Second, the fact of the matter is that non-price competition effects, namely quality, reliability, marketing strategy and etc., are generally excluded from standard export demand equations. We feel that even such a crude commodity composition index may give some clues considering non-price effects such as product innovation process. Following Krugman (1989), a measure of the "range of goods" traded in export markets may be included in the foreign trade regressions to capture the supply effects. We believe that our exports commodity composition index, \( XCC \), is a proxy for the range of goods traded in the export markets, and thus captures the supply effects in the Krugman's sense.

5. Econometric Methodology

Cointegration and Granger causality between variables and the short-run dynamic adjustment towards the long-run equilibrium path is to be examined. The appeal of the cointegration analysis for economists is that it simply provides a formal framework for testing and modeling long-run economic relationships from actual time series data.

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\(^4\) For various measures of competitiveness, see Shone (1989).  
\(^5\) For a detailed description of the variables employed in this study and the data sources, see Appendix.  
\(^6\) Economic theory provides insight into how each variable in Equations (4) and (5) should affect exports and imports. As regards the export demand equation, the higher the level of foreign real income activity, \( ceteris paribus \), the larger would be foreign demand for the country's exports. The higher the price of the country's exports relative to those of other countries, \( ceteris paribus \), the smaller would be the demand for the country's exports. The similar logic applies to the variables in the import demand equations.  
\(^7\) For the same point, see Muscatelli et al. (1991).  
\(^8\) In examining the empirical relevance of Krugman's supply effects, Madsen and Damania (1994) employes the supply of manufactures as an instrument for the range of goods.
This involves the 'two-step procedure' suggested by Engle and Granger (1987) (EG hereafter). As a first step, we estimate the following cointegrating regressions by ordinary least squares (OLS):

\[ X_t = \alpha_0 + \beta_0 Y_t + \mu_t \]  
\[ Y_t = \alpha_1 + \beta_1 X_t + \mu'_t \]  

(6)  
(7)

where \( \alpha_0 \) and \( \alpha_1 \) represent the intercept terms while \( \mu_t \) and \( \mu'_t \) are the error terms.

First we check for the cointegrating properties of the series involved. The next step is to test which variable Granger causes the other,\(^9\) using error-correction models (ECMs) to see if the coefficient of the error-correction term is statistically significant or not.\(^11\) Accordingly, the ECMs are formulated as follows:

\[ \Delta X_t = a_0 - b_0 \mu_{t-1} + \sum_{i=1}^{m} c_i \Delta X_{t-i} + \sum_{j=1}^{n} d_i \Delta Y_{t-j} + e_t \]  
\[ \Delta Y_t = a_1 - b_1 \mu'_{t-1} + \sum_{i=1}^{q} e_i \Delta Y_{t-i} + \sum_{j=1}^{r} f_i \Delta X_{t-j} + e'_t \]  

(8)  
(9)

where \( \mu_{t-1} \) and \( \mu'_{t-1} \) are the lagged estimated residuals (i.e. error-correction terms) derived from the static cointegrating regressions (6) and (7) respectively. The term \( \Delta \) represents the first differences. Statistically significant \( b_0 \) and \( b_1 \) suggest that \( Y \) Granger causes \( X \) and \( X \) Granger causes \( Y \) respectively.\(^12\) The ECMs introduce an additional channel through which Granger-causality could be detected since if two variables are cointegrated, causality must run in, at least, one direction between them. This causal relationship between the two variables provides the short-run dynamics necessary to obtain long-run equilibrium (Granger, 1988). For instance, focusing on equation (8), \( Y \) is said to Granger cause \( X \) not only if the \( d_i \)'s are jointly significant, but also if \( b_0 \) is significant. Thus, in contrast to the standard Granger test for causality, the ECMs allow for the finding that \( Y \) Granger causes \( X \), as long as the error-correction term, \( \mu_{t-1} \), carries a significant coefficient even if the \( d_i \)'s are not jointly significant (Granger, 1988). Jones and Joulfaian (1991) support the interpretation that the changes in the lagged independent variable describes the short-run causal impact, while the error-correction term introduces the long-run effect. However, if the two variables are not cointegrated, then the error-correction terms are dropped from the ECMs and the standard Granger test for causality is carried out.

We apply the integration and cointegration analyses in the EG sense; that is, a time series, say, \( X_t \) is said to be integrated of order \( d \) if, after differencing \( d \) times, it becomes stationary, denoted as \( X \sim I(d) \). Moreover, two time series, \( X_t \) and \( Y_t \) are said to be cointegrated of order \( d, b \) where \( d \geq b \geq 0 \), denoted as \( X_t, Y_t \sim CI(d, b) \) if:

\begin{enumerate}
  \item[a)] both are \( I(d) \), and \( b) \) their linear combination \( \alpha_j X_t + \alpha_2 Y_t \) is \( I(d - b) \); that is, the residuals of the long-run regression should be stationary (i.e. integrated of order zero). The vector \( [\alpha_j, \alpha_2] \) is referred to as the 'cointegrating vector'.
\end{enumerate}

In the present paper, we empirically investigate the multivariate version of the relationship stated, i.e. single equation multivariate cointegration analysis. Here, we mainly rely on the econometric methodology of

---

\(^9\) As OLS is super consistent in the cointegrating regressions, asymptotically it is not relevant whether these regressions are normalised on \( X \) or \( Y \). In finite sample, however, the normalisation may matter, and we consider both possibilities.

\(^10\) In this paper, we use causality in Granger's sense.

\(^11\) This is known as the Granger Representation Theorem (GRT). See Engle and Granger (1987). According to the GRT, if two time series are cointegrated, then there exists an error-correction representation (i.e. error-correction mechanism is well determined) and vice versa. Note that in small samples, statistically significant estimates of \( b \) in equations (8) and (9), provide further evidence that the variables in eq. (6) and (7) are indeed cointegrated.

\(^12\) Note that joint significance of the error-correction terms \( b_0 \) and \( b_1 \) could be a matter of debate.
the EG outlined above with some necessary corrections / modifications to deal with the endogeneity problem, i.e. fully modified unrestricted ECM (Inder, 1993). One drawback of the EG type of single equation modelling is that it assumes uniqueness of the cointegrating vector. However, in a multivariate context the number of cointegrating vectors could be more than one (i.e. r > 1). If r > 1, there is no longer a unique long-run relationship towards which the error-correction model (ECM) is adjusting. In this case, a single equation cointegrating regression will estimate the linear combination of the existent vectors. Although the existence of multiple cointegrating vectors is seen as an identification problem, applied researchers overcome this problem by choosing the cointegrating vectors which makes 'economic sense'. This implies choosing the cointegrating vector where the estimated long-run elasticities correspond closely (in both magnitude and sign) to those predicted by economic theory.

The validity of conditional models relies on the exogeneity of the variables on which we condition. Alternatively, if they cannot be treated as weakly exogenous, then one should use the appropriate correction mechanism to tackle the endogeneity bias. A number of tests for weak exogeneity in cointegrated variables have been proposed in recent years (for an evaluation of these tests and definitions, see Urbain, 1993). Among them we follow the EG (1987). Within their two-step framework, EG argues that a simple way to check the weak exogeneity of, say, explanatory variable X_t for the long-run and short-run parameters of interest is to estimate an ECM for X_t and test the statistical significance of the error-correction term using a traditional t-test. If the t-test is significant, then X_t can no longer be treated as weakly exogenous.

6. Empirical Findings

Data

In the light of the econometric methodology developed in the earlier section, we now apply the single equation multivariate cointegration analysis and the ECMs to examine the import and export demand equations between Turkey and the EU. We use annual data for the period 1963-2002 for our single equation multivariate cointegration analysis with ECM. The variables used in our import demand cointegrating regression are the MV, RPM, YD and D. They simply correspond to import volume of Turkey from the EU, relative import prices of Turkey regarding the EU, GDP volume index of Turkey, and the Turkish external debt stock respectively. As regards the export demand cointegrating regression, the variables included are the Turkish export volume towards the EU, relative export prices (three different proxies are used) regarding the EU, GDP volume of the export market, i.e. the EU (two different proxies are used). The variables used in our export demand cointegrating regression are the XV, RPX1, RPX2, RPX3, MVEU, YEU and XCC. We also employ two dummies to capture the effects of the possible structural breaks in the import and export demand cointegrating regressions, namely DU1977, DU1987. This is due to the fact that breaks may result in supurious roots regarding residual-based cointegration tests. We use the natural logarithm of the relevant variables, except variable XCC, (prefixed with the letter L), since their first differences reflect the rate of change of each variable. Data definitions, data sources and further information are provided in the Appendix.

Integration (Dickey-Fuller) and cointegration (Engle-Granger) analyses: standard approach

In the light of the theory and the econometric methodology outlined, we first examine the multivariate cointegration and causality issues among the import and export demand variables considered. We are mainly interested in analysing the following classical multivariate import and export demand relationships (same with eq. 4 and 5):

\[ MV = f(\text{RPM}, YD) \]  \hspace{1cm} (10)
\[ XV = f(\text{RPX}, YEU) \]  \hspace{1cm} (11)

13 It is important to note that the choice between system-based models and conditional (single-equation) models is not straightforward, and is also open to debate. Urbain (1993) points out that if some exogeneity conditions are satisfied, a single equation models, from a practical point of view, enjoy nice asymptotic properties.

14 The standard orthogonality tests (such as the Hausman test) in the presence of cointegrated variables may well be invalid due to nonstationary nature of the variables in levels, and the null hypothesis is usually not sufficient for weak exogeneity in cointegrated models (for this point, see Urbain, 1993).
Following the methodology for multivariate analysis set up in the earlier section, we now express this longrun relationship as a regression in natural logarithms:

\[
LMV_t = \alpha_0 + \alpha_1 LRPM_t + \alpha_2 LYD_t + \mu_t \\
LXV_t = \beta_0 + \beta_1 LRPX_t + \beta_2 LYEU_t + \mu' \tag{12}
\]

where \( \mu_t \) and \( \mu' \) are the residuals. For the variables in (12 and 13) to be cointegrated, they need to be I(1) as a necessary condition (but not sufficient). Table 3 suggest that all variables employed in our long-run regressions are integrated of order one, i.e. their first differences are stationary.

### Table 3: The ADF test for integration level

<table>
<thead>
<tr>
<th>Variables</th>
<th>Test Statistic</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>levels first differences</td>
</tr>
<tr>
<td><strong>Import Demand Equation</strong></td>
<td></td>
</tr>
<tr>
<td>LMV</td>
<td>-1.98(0)</td>
</tr>
<tr>
<td>LRPM</td>
<td>-2.00(1)</td>
</tr>
<tr>
<td>LYD</td>
<td>-2.20(0)</td>
</tr>
<tr>
<td>LD</td>
<td>-1.81(1)</td>
</tr>
<tr>
<td>DU1977</td>
<td>-2.02(1)</td>
</tr>
<tr>
<td><strong>Export Demand Equation</strong></td>
<td></td>
</tr>
<tr>
<td>LXV</td>
<td>-2.61(0)</td>
</tr>
<tr>
<td>LRPX1</td>
<td>-1.93(1)</td>
</tr>
<tr>
<td>LRPX2</td>
<td>-2.20(1)</td>
</tr>
<tr>
<td>LRPX3</td>
<td>-2.77(1)</td>
</tr>
<tr>
<td>LMVEU</td>
<td>-1.89(1)</td>
</tr>
<tr>
<td>LYEU</td>
<td>-2.32(1)</td>
</tr>
<tr>
<td>XCC</td>
<td>-2.23(2)</td>
</tr>
<tr>
<td>DU1987</td>
<td>-1.95(1)</td>
</tr>
</tbody>
</table>

**Note:** Intercept term included in the ADF equations. Time trend is included in the ADF equations only if statistically significant. The corresponding critical values with and without time trend (obtained from MacKinnon, 1991) for 5% significance level are –3.53 and –2.94 respectively. Figures in parentheses show the number of augmentation that sufficient to secure lack of autocorrelation of the error terms.

A sufficient condition for a joint cointegration among the variables now is that the error term (\( \mu_t \)) of the cointegrating regression should be stationary. The residual-based ADF test statistic for \( \mu_t \) suggest that we cannot reject the null of no cointegration at 5 per cent significance level (see equation 5). Following is the estimation resultsof the EG cointegrating regression (12) by OLS.

**Import Demand: long-run**

\[
LMV_t = -8.55 - 0.99 LRPM_t + 2.54 LYD_t + \mu_t \\
( -11.0)(-6.37) (16.4)
\]

\( R^2 = 0.94 \) \( \text{RSS} = 1.79 \) \( \text{CRDW} = 0.78 \)

ADF = -3.01 (corresponding critical value at 5% is –3.97)

Sample: annual data (1963-2002) t-statistics are reported in parantheses, having only a descriptive role since the variables are non stationary.

---

It is important to note that we reestimates the long-run export regression (15) by replacing RPX1 with RPX2 and RPX3. Results with RPX3 produces similar ones. However, using RPX2 as a measure for relative export prices yields insignificant price variable in which Turkish firms are assumed to be competing with their Asian competitor in the EU market. As regards the scale variable in export equation (15), we obtained similar estimate results when replace MVEU with YEU (available on request).
**Export Demand: long-run**

\[ L_{XV_t} = -0.20 - 2.52L_{RPX_t} + 1.96L_{MVEU_t} + \mu_t, \]  
\[ (-0.91)(-9.87) \quad (17.8) \]  

\( R^2 = 0.96 \quad \text{RSS} = 1.88 \quad \text{CRDW} = 0.66 \)  

\( \text{ADF} = -2.93 \quad \text{(corresponding critical value at 5% is –3.97)} \)

Note again that the estimated t-statistics and other standard test statistics in (14) and (15) have only a descriptive role since the variables are non stationary (Banerjee et al., 1986). Since the residual-based ADF test statistics –3.01 and –2.93 are smaller than the corresponding critical value –3.97 at 5% statistical significance level, we cannot reject the null of no joint cointegration among the variables against the alternative. Since \( R^2 > \text{CRDW} \), the joint cointegration cannot be ensured (see Banerjee et al., 1986). This result may well be due to the low power of cointegrating statistics of the residual-based type. One possibility is that structural breaks/changes may result in spurious unit roots, i.e. failure of rejection of the null hypothesis. Another possibility is that the omission of some relevant variables in the long-run regression may have caused the above results. In what follows, we analyse the effects of breaks on the cointegration tests and the inclusion of some relevant variables in the regressions.

**Unit Roots / Cointegration with breaks and the Supply / Product Innovation Effects**

The results in Table 3 suggest that all variables appear to be stationary in first differences, i.e. I(1). These results are validated despite some structural breaks/changes. Our Perron unit root test results are available on request (for the method see Perron 1990; Perron and Vogelsang, 1992). In order to see the effect of the structural changes and/or regime shifts on the cointegrating regressions we take two steps. First, we include dummies for the possible break years and other relevant variables in the static import and export demand cointegrating regressions. Second, we test for cointegration with breaks using the methodology suggested by Gregory and Hansen (1996). This methodology examines the presence of cointegrated relationship under possible regime-shifts and use suggest three different models.

**Import Demand: long-run**

\[ L_{MV_t} = -9.39 - 0.37L_{RPM_t} + 3.33L_{YD_t} - 0.25L_{D_t} - 0.56L_{U1977} + \mu_t, \]  
\[ (-19.6)(-2.94) \quad (13.6) \quad (-2.58) \quad (-5.37) \]  

\( R^2 = 0.98 \quad \text{RSS} = 0.54 \quad \text{CRDW} = 1.60 \)  

\( \text{ADF} = -5.14 \quad \text{(corresponding critical values at 5% is –4.80)} \)

Sample: annual data (1963-2002) t-statistics are reported in parantheses, having only a descriptive role since the variables are non stationary.

**Export Demand: long-run**

\[ L_{XV_t} = 0.70 - 1.19L_{RPX_t} + 0.90L_{MVEU_t} + 3.95X_{CC_t} + 0.16L_{U1987} + \mu_t, \]  
\[ (5.43) \quad (-7.21) \quad (8.39) \quad (8.53) \quad (2.31) \]  

\( R^2 = 0.99 \quad \text{RSS} = 0.38 \quad \text{CRDW} = 1.78 \)  

\( \text{ADF} = -5.49 \quad \text{(corresponding critical value at 5% is –4.79)} \)

Sample: annual data (1963-2002) t-statistics are reported in parantheses, having only a descriptive role since the variables are non stationary.

We can now reject the null of no cointegration for cointegrating regressions (16) and (17) at 5% and/or 10% significance levels. As regard the long-run import demand regression (16), two new variables are included, namely LD and DU1977. They stand for the natural logarithm of external debt stock of Turkey and the dummy variable for the possible break year 1977 respectively. They both produced reasonably high t-
statistics. In addition to the external debt stock variable, we used some other measures for ‘import capacity’ such as the indicator of foreign exchange availability (INFLOW) proxied by exports earnings and non-gold international reserves (RES). However, unlike the external debt stock variable D, the variables INFLOW and RES in natural logarithms produced very low t-statistics, and thus not included in the final model.

It is clear from eq. (16) that the inclusion of the debt variable as a measure for import capacity and the dummy for the break year worked out well. There exists a joint cointegration among the variables according to residual-based ADF cointegration tests. Banerjee et al. (1986) propose simple and quick rule; that is, if $CRDW > R^2$, the null of no cointegration is more likely to be rejected. For both eq. 16 and 17, we have $CRDW > R^2$. There is also strong evidence in eq. 16 that, the inclusion of variables LD and DU1977 produced even lower elasticity for relative import prices, -0.32, and higher elasticity for income variable. For both long-run regressions 16 and 17, we employed Engle-Yoo corrections (see Engle and Yoo, 1991) and all variables produced statistically significant t-statistics (available on request). Finally, we have inelastic significant relative import prices, and very elastic income variable as regards the long-run cointegrating import demand regression.

As far as the export demand equation is concerned, the large scale and price elasticities may represent hidden structural breaks, due to substantial change in commodity composition towards high-technology products, product innovation and/or process innovation. In order to see the effects of these on export volume, we developed an index, XCC, which captures the increasing exports of manufactures of LDCs, especially those Newly Industrialising Countries' (NICs), in what is regarded high-technology products, whose demand tends to be income and perhaps price elastic. Such an index may then be introduced into the export demand equation to see if allowing for commodity composition, even in such a simple way, has a significant effect on the estimated scale and price elasticities of export demand. Accordingly, one would expect export demand to be positively correlated with this commodity composition index. XCC might also be seen as a proxy for the range of goods traded in the export markets of Turkey, and thus captures the supply effects in the Krugman's sense. Krugman (1989) suggests that the growth process of the NIC’s has been driven by a continous process of product innovation and diversification.

The index, XCC, is constructed in the following way: first, export goods is divided into four commodity groups, $(C_1,...C_4)$. These groups are selected in such a way as to include products with increasing technological content, as we move from $C_1$ to $C_4$. The second stage is to construct an index, XCC, which lies over interval $[0,1]$. Regarding the export equation (17), as expected, inclusion of the export commodity composition index, XCC, in log terms resulted in smaller income and price elasticities. This is so because the commodity type effects are implicitly captured by the income and price effects if they are not represented in the regression. As we understand from equation (17), export commodity composition change from agricultural to manufacturing / technology intensified products can statistically explain the increase in Turkish export volume towards the EU. This also shows the significant effect of product / process innovation processes undertaken in Turkey. Equation (17) produced price and income elasticities close to unity.

We also test for cointegration with breaks using the methodology suggested by Gregory and Hansen (1996). This methodology examines the presence of cointegrated relationship under possible regime-shifts and use suggest three different models. In this paper we prefer the model 3, i.e. regime shift (C/S) (see Gregory and Hansen, 1996, 103). Using the model with regime shift (C/S) one gets some tests statistics including ADF\*$. The corresponding critical values are obtained from Gregory and Hansen (1996, 109). Regarding our analysis, the ADF\* test statistic provides empirical support for the presence of cointegration with possible breaks among the variables concerned (available on request).

**Unique cointegrating vector (Johansen, VAR)**

We now test if this is the only cointegrating vector or not by applying the Johansen ML VAR test

\[16\] For details of the definition of XCC see Appendix.
procedure (Johansen, 1988). Our results confirm the unique cointegrating vector for both export and import models. Table 4 shows the results.

Table 4. Johansen ML test for cointegration (Maximum Eigenvalue Test / VAR=1): Import and Export Models

<table>
<thead>
<tr>
<th></th>
<th>Import Demand Model</th>
<th></th>
<th></th>
<th></th>
<th></th>
<th>Export Demand Model</th>
<th></th>
<th></th>
<th></th>
<th></th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Null</td>
<td>Alternative</td>
<td>Test Statistic</td>
<td>Critical Value</td>
<td></td>
<td>Null</td>
<td>Alternative</td>
<td>Test Statistic</td>
<td>Critical Value</td>
<td></td>
</tr>
<tr>
<td></td>
<td>r(=0)</td>
<td>r(=1)</td>
<td>40.18</td>
<td>28.72</td>
<td></td>
<td>r(\leq 1)</td>
<td>r(=2)</td>
<td>14.58</td>
<td>22.16</td>
<td></td>
</tr>
<tr>
<td></td>
<td>r(\leq 1)</td>
<td>r(=2)</td>
<td>14.58</td>
<td>22.16</td>
<td></td>
<td>r(\leq 2)</td>
<td>r(=3)</td>
<td>8.01</td>
<td>15.44</td>
<td></td>
</tr>
<tr>
<td></td>
<td>r(\leq 2)</td>
<td>r(=3)</td>
<td></td>
<td></td>
<td></td>
<td>r(\leq 3)</td>
<td>r(=4)</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td></td>
<td>r(\leq 3)</td>
<td>r(=4)</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

Note: Critical values are obtained from MacKinnon (1991).

Endogeneity bias and comparison of different approaches

However, the long-run OLS estimators are biased if the explanatory variables are not weakly exogenous. Only if they are weakly exogenous, we can assume away the 'endogeneity bias'. If not, an appropriate correction for OLS estimators will be necessary. As mentioned earlier, EG argue that a simple way to check the weak exogeneity of, say, explanatory variable \(X_t\) for the long-run and short-run parameters of interest is to estimate an ECM for \(X_t\) and test the statistical significance of the error-correction term using a traditional t-test. If the t-statistics is significant, then \(X_t\) can no longer be treated as weakly exogenous.

Our calculations show that as regards the import model \(LRPM\) and \(LD\) are not weakly exogenous while all explanatory variables are endogenous and cannot be treated weakly exogenous. Accordingly, we apply the fully modified ECM method to get the long-run estimators which are free from 'endogeneity' bias. Employing the methodology suggested by Inder (1993), we get the fully modified unrestricted ECM estimates. Table 5 reports the long-run estimates obtained by using different approaches. Results reported in Table 5 suggest that our long-run estimates are quite robust. For better comparison, we also calculate the fully modified Phillips-Hansen estimator, free from nuisance parameter effects (Phillips and Hansen, 1990), and the asymptotically efficient dynamic OLS estimates of Saikkonen (Saikkonen, 1991).

Table 5: Estimates of our long-run relationship: a comparison of different approaches

<table>
<thead>
<tr>
<th>Variable</th>
<th>Static EG OLS (Engle&amp;Granger)</th>
<th>Fully Mod. Unr. ECM (Inder)</th>
<th>Dyn. OLS (Saikkonen)</th>
<th>Fully Mod.OLS Phillips&amp;Hansen</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>LRPM</td>
<td>-0.37</td>
<td>-0.26</td>
<td>-0.56</td>
</tr>
<tr>
<td></td>
<td>LYD</td>
<td>3.33</td>
<td>3.21</td>
<td>3.76</td>
</tr>
<tr>
<td></td>
<td>LD</td>
<td>-0.25</td>
<td>-0.27</td>
<td>-0.38</td>
</tr>
<tr>
<td>Export Demand Model</td>
<td>LRX1</td>
<td>-1.19</td>
<td>-1.28</td>
<td>-0.91</td>
</tr>
<tr>
<td></td>
<td>LMVEU</td>
<td>0.90</td>
<td>0.92</td>
<td>0.77</td>
</tr>
<tr>
<td></td>
<td>XCC</td>
<td>3.95</td>
<td>5.61</td>
<td>5.19</td>
</tr>
</tbody>
</table>

Note: Own estimates.

It is important to point out that long-run estimates reported in Tables 5 are free from possible endogeneity bias. Note that long-run estimates by different methods are quite robust, and well compatible with the EG static estimates reported earlier.

The EU Integration Effect on the Trade

Following Krugman (1989) it is suggested that if a measure of the range of goods traded in international markets is included in the trade regressions (which we have already done earlier in this paper by
including the variable XCC in the export model), together with a proxy for the level of integration in international markets, then the income elasticity declines. This implies that the high income elasticities found in previous works may in fact result from the omission of integration and supply effects (for an exception see Madsen and Damania, 1994).

So far we made no attempt to measure the impact of integration of markets on foreign trade regarding the integration of the markets of the EU and Turkey over the period 1963-2003. A simple way in which integration factor can be incorporated is by including a time trend in the model. This naturally presumes that the process of integration has been continuous. Thus we include a time trend in the model to allow for the possibility of continuous integration effects. It can of course be argued that the time trend may be measuring some other variables that have been omitted from the cointegrating model. We admit that in the absence of more accurate measure of integration, the precise interpretation of this variable remains ambiguous.

Our empirical results for the import demand model show that time trend is insignificant and the long-run parameter estimates change only slightly when integration effects are allowed for. However, results for export model produce high t-statistics for the time trend. It is also clear that the long-run parameter estimates for the export model change significantly when integration effects are included. This implies that Turkish export demand in the EU market has been significantly affected by the integration process. The following equation presents the results for export model (T for time trend):

\[
LXV_t = 0.32 - 0.97LRPX_t + 0.61LMVEU_t + 3.05XCC_t + 0.13DU1987 + 0.03T + \mu_t',
\]

\((17)\)

\( \begin{array}{cccccc}
(1.66) & (-5.44) & (4.03) & (5.45) & (1.88) & (2.53)
\end{array} \)

\(R^2 = 0.99\) \quad \text{RSS} = 0.32 \quad \text{CRDW} = 1.92

Sample: annual data (1963-2002) t-statistics are reported in parentheses, having only a descriptive role since the variables are non stationary

The results in equation (17) reveal that the long-run parameter estimates change remarkably when integration effects are allowed for. The long-run income elasticity is 0.61, and well below the estimates in the previous long-run versions. The parameter estimate for relative prices is 0.97 and also below the estimates in the previous parts.

_error Correction Models, ECM / Short-run and Causality_

To show the multivariate causal effect, we now apply the Granger causality test. Since, after all, EG OLS estimates were shown to be robust, the estimated lagged residuals may still be used in the ECM as the error-correction term. Table 6 and 7 show the Granger causality test results from the ECMs.

| Table 6: Import Demand: Error correction model/Short-run/Granger causality: multivariate case |
|---------------------------------|-----------------|-----------------|-----------------|-----------------|-----------------|
| Dependent Variable             | t-statistic     | F-Statistic     | F-Statistic     | F-Statistic     | F-Statistic     |
| for \(\mu_{t-1}\)             | for \(\Sigma \Delta LMV\) | for \(\Sigma \Delta RPM\) | for \(\Sigma \Delta GDPV\) | for \(\Sigma \Delta LD\) |
| \(\Delta LMV\)               | -0.62*         | 6.54(2)*       | N.S.            | 29.9(1)*       | N.S.            |

\(\text{Note:} \quad \mu_{t-1}\) denotes the error correction term. Numbers in parantheses indicate the number of lags. Note that optimum number of lags are determined by applying general-to-specific methodology. \(\Delta\) represent first differences.

\(\text{* significant at } 1\% \quad \text{N.S. Not significant}\)
### Table 7: Export Demand: Error correction model / Short-run / Granger causality: multivariate case

<table>
<thead>
<tr>
<th>Dependent Variable</th>
<th>for $\mu_{t-1}$</th>
<th>for $\Sigma \Delta LXV$</th>
<th>for $\Sigma \Delta LRPX1$</th>
<th>for $\Sigma \Delta LMVEU$</th>
<th>for $\Sigma \Delta XCC$</th>
</tr>
</thead>
<tbody>
<tr>
<td>$\Delta LXV$</td>
<td>-0.71</td>
<td>N.S.</td>
<td>36.4(0)</td>
<td>23.7(1)</td>
<td>23.9(0)</td>
</tr>
</tbody>
</table>

*Note: $\mu_{t-1}$ denotes the error correction term. Numbers in parentheses indicate the number of lags. Note that optimum number of lags are determined by applying general-to-specific methodology. $\Delta$ represent first differences. $^*$ significant at 1% -- N.S. Not significant*

We have evidence that explanatory variables Granger cause dependent variables $LXV$ and $LMV$ through two channels: *first*, they jointly Granger cause the dependent variables through the statistically significant error correction terms and *second*, some variables have Granger cause effect separately (see the joint significance F-statistics in Table 6 and 7). We have the long-run causal effect via the first one while the second causal effect has a short-run character (Jones and Joulfaian, 1991).

### 7. Conclusion

On the basis of the cointegration analyses employed, we observe genuine long-run relationships among the statistically significant variables regarding the export and import cointegrating regressions. Our findings suggest that conventional estimates of import and export equations which estimate long-run elasticities in excess of unity may reflect omitted variable bias, and may represent hidden structural breaks. This paper deals with the possible effects of factors such as structural breaks, integration of markets, product innovation, supply, and omitted variables as regards the significance and the magnitude of the income and price elasticities. The various long-run estimates obtained here reveal that the inclusion of a supply variable (i.e. commodity composition index, $XCC$), dummies for structural breaks, and a measure of economic integration with the EU significantly lowers both the long-run price and income elasticities for Turkish exports with the EU. Regardind the import function, the inclusion of dummies for structural break and a measure of import capacity (i.e. external debt stock) lowers the price elasticity although the income elasticity remained high and significant. The corresponding short-run error correction models seem to be well specified and working well.

As regards the debate of conventional wisdom (*elasticity pessimism*) versus alternative paradigm (*elasticity optimism*) on the size and the significance of the income and price elasticities for export demand, we provide some econometric evidence suggesting that both income and price effects have been the driving forces. The Marshall-Lerner condition is moderately satisfied $[(-0.37) + (-0.97) = -1.34]$. However, the income elasticity for the import demand equation is significant and very elastic (3.33) so that this might have offsetting effects on the exchange rate adjustments. Overall, estimation results support the view that the success of Turkish exports in the EU export market cannot only be attributed to high level of devaluation occurred especially during the 1980s. Non-price factors such as the exports commodity composition index, $XCC$, even in such a simple and crude fashion, are shown to be significant to explain the successful export drive of Turkey. When evaluated in the Krugman's sense, our $XCC$ implies significant supply effects in assessing the success story of Turkish export performance. It is also important to note that a recession or stagflationary and protectionist policies in the export markets can easily lead to substantial reductions in export demand, *ceteris paribus*, if world income is a statistically significant factor. Our long-run Turkish export estimates for income variable yields low but statistically significant income elasticity regarding the EU.

### Appendix

#### Data Sources

The data used in this study are annual for the period of 1963-2002 and are taken from the following sources: IMF, International Financial Statistics; CBRT (Central Bank of Turkish Republic); SPO (State Planning Organization / Turkey); SIS (State Institute of Statistics / Turkey).
Definitions of the Variables

**XV:** Exports of goods, volume index (1980=100) constructed on the basis of the following formula:

\[ XV = \frac{X$}{PX} \]

where \( X$ \) and \( PX \) represent total exports to the main six traditionally trading countries within the EU, namely Germany (GER), France (FRA), Italy (ITA), UK, Nederland (NED) and Belgium (BEL) (which forms more than 50% of the total Turkish exports to the EU) in US dollars, and export price index in US dollar terms respectively.

**RPX:** Relative export prices, that is the ratio of export prices of Turkey to the prices that Turkish companies face in the export market (i.e. world prices), both expressed in US dollar terms. It can also be referred to as "export price competitiveness" or "real exchange rate" of the country (for same sort of real exchange rate definition and various measures of competitiveness, see e.g. Shone, 1989). We use three different measures for RPX: RPX1, RPX2 and RPX3.

\[
\begin{align*}
\text{RPX1} &= \frac{PX}{PMEU} \\
\text{RPX2} &= \frac{PX}{PXASIA} \\
\text{RPX3} &= \frac{PX}{TUFEEU}
\end{align*}
\]

where

- \( PX \) is the US dollar-based export price index of Turkey (1980=100).
- \( PMEU \) is the US dollar based-import price index of the EU countries of UK, GER, ITA, NED (average). (1980=100)
- \( PXASIA \) represents the US dollar based-export price index of the Asian countries (average) (1980=100)
- \( TUFEEU \) stands for the consumer price index of the EU countries of UK, FRA, GER, ITA, NED, BEL (average). (1980=100)

**YEU:** The real income (GDP) expressed as an index (1980=100) in the EU export market facing Turkey. GDP volume index of the EU countries of UK, FRA, GER, ITA, NED, BEL (average) (1980=100)

\[ \text{MVEU} = \frac{MEU}{PMEU} \]

MVEU is the total real imports of the EU countries UK, GER, ITA, NED expressed as volume index (1980=100) where MEU and PMEU represent total imports of the EU countries UK, GER, ITA, NED in US dollars, and average import price index of the EU countries UK, GER, ITA, NED in US dollar terms respectively.

\[ \text{RPM} = \frac{PM}{PD} \]

RPM is the relative import prices, that is the ratio of import prices of the EU facing Turkey (\( PM \)) to domestic prices (wholesale price index) (\( PD \)), both expressed in US dollar terms. It can also be referred to as "import price competitiveness" where

**YD:** GDP volume index of Turkey (1980=100).

**RES:** The ratio of non-gold international reserves of Turkey expressed in US dollar.

**INFLOW:** It represents foreign exchange inflows to Turkey. Annual earnings of total exports in US dollars.

**D:** External debt stock of Turkey expressed in US dollars.

\[ \text{DU1977: } DU_{1977} = 1 \text{ if } t>T_{1977} \text{ and 0 otherwise.} \]

\[ \text{DU1987: } DU_{1987} = 1 \text{ if } t>T_{1987} \text{ and 0 otherwise.} \]
The commodity composition index, XCC, is constructed in the following way: first, export goods is divided into four commodity groups, \( (C_1 \ldots C_4) \). These groups are selected in such a way as to include products with increasing technological content, as we move from \( C_1 \) to \( C_4 \). The second stage is to construct an index, \( XCC_t \), which lies over interval \([0,1]\). We follow Muscatelli et al. (1991 and Utkulu (1995) in constructing the index, and in choosing a symmetric distribution for the weights, i.e. the weights chosen are: \( a_1 = 0, a_2 = 0.33, a_3 = 0.67, a_4 = 1 \), over the interval \([0,1]\):

\[
XCC_t = \frac{\sum_{t=1}^{4} a_t C_t}{\sum_{t=1}^{4} C_t}
\]

**Commodity Groups for the Index, XCC** (the difference between the two versions of SITC, namely SITC Revised 2 and SITC Revised, has also been taken into account by converting SITC Revised into SITC Revised 2):

- **\( C_1 \)**: Total exports of agricultural products and crude materials, (SITC Revised 2 groups, 0 and 2).
- **\( C_2 \)**: Total exports of traditional manufacturing sectors, (SITC Revised 2 groups, 61, 62, 63, 64, 65, 69, 84, 85, 89).
- **\( C_4 \)**: Total exports of specialised supply and science-based sectors, (SITC Revised 2 groups, 71, 72, 73, 74, 75, 87, 88).

**References**


