Price and Income Elasticities of Turkish Export Demand: A Panel Data Application

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Abstract

In this paper, price and income elasticities of export demand are calculated. The study is extended to sectoral and country specific export demand functions. The paper presents some panel unit root and cointegration tests, which have been studied extensively in recent years. The major aim of this study is to find the price and foreign income elasticities of aggregate export demand. According to the estimation results, the real exchange rate elasticity of total export demand is found to be less than one, whereas the income elasticity is found to be greater than one.

Keywords: Panel Unit Root Test, Panel Cointegration Test, Income and Real Exchange Rate Elasticities
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1. Introduction

Export is generally considered to play an important role in the economic development of a country. In this respect, measuring the income and price elasticities of foreign trade, especially in developing countries, has received a great deal of attention because of its substantial implications on trade policy and balance of payments issues. Senhadji and Montenegro (1999) emphasized the importance of export demand elasticities as follows; demand elasticity is a measure of sensitivity of demand against the changes in price and income. The higher the income elasticity of export demand, the more powerful exports will be as an engine of growth. The higher the price elasticity, the more competitive is the international market for exports of the particular country, and thus a real devaluation will be more successful in promoting the export revenues. Accordingly, price and income elasticities of export demand become important for investigating the effects of devaluation on trade balance. Based on this statement, the major aim of this study is to find the price and foreign income elasticities of aggregate export demand.

Turkey's export performance has been studied widely in recent years, most of which concentrated on the relationship between export growth and economic growth. For example, Arslan and van Wijnbergen (1993) investigate the driving forces of Turkey's export boom during the years 1980 and 1987. The main interest of the relevant study is the relative contributions of export subsidies and real exchange rate depreciation to Turkey's export growth. With this aim, export supply and demand equations are estimated. The export supply equation shows the response of export supply to relative prices (export prices and home prices of goods) inclusive of export subsidies. The export demand equation is estimated separately for oil exporting and non-oil exporting countries. The relative price elasticity and foreign income elasticity of export demand are found to be significant and high in both equations. After simulation analysis they concluded that the policies allowing real depreciation of the exchange rate were more effective than export incentives on export growth. Bahmani-Oskooee and Domac (1995) investigated the export-led growth hypothesis for Turkey using cointegration analysis with annual data covering a wide time span from 1923 to 1990. They concluded that there is a long-run equilibrium relationship between export growth and output growth in Turkey. Their findings show that there is also a bi-directional Granger causality between the variables, i.e., not only export growth causes output growth, but also output growth causes export growth. Conversly, Özmen and Furtun

(1998) rejected the validity of the export-led growth hypothesis for Turkey. On the other hand, Yiğidim and Köse (1997) investigated the validity of export-led growth hypothesis for Turkey by constructing a regression equation where GDP growth being the dependent variable and import and export being the explanatory variables. Their findings indicate that export is not statistically significant on the economic growth, yet import is the most important determinant of the economic growth in Turkey.

There are also some studies, which are focused especially on the effects of exchange rate on trade volume. Bahmani-Oskoode and Ltaifa (1992) estimated a real export equation using cross-sectional data in which the explanatory variables were rate of devaluation of each country's exchange rate against the US dollar, real income, population and measure of exchange rate variability of country i. Their results indicate that, for Turkey, the coefficient of exchange rate is significant and has negative effect on export. Sivri and Usta (2001) investigated the relationship between the real exchange rate, export and import by using the VAR model. According to variance decomposition analysis, the real exchange rate does not explain an important proportion of the forecast variance of exports and imports. In addition to this, it Granger causes neither exports nor imports. Sivri and Usta concluded that the real exchange rate couldn't be used as a trade policy instrument effectively.

One of the studies investigating the Turkish export boom in the 1980s is the study of Barlow and Şenses (1995). They attempted to measure the extent to which export growth was due to the policies undertaken or due to the external circumstances such as the Iran-Iraq war, changes in consumer income of Turkey's major trading partners and rainfall fluctuations. They estimated export supply and demand equations on disaggregated level, i.e. export of agricultural and manufactured commodities and foreign demand of industrial oil-producing countries. Their results can be summarized as follows; income of industrial countries is a significant predictor of both agricultural and manufactured exports. The real exchange rate, the export subsidy rate and the level of income in the oil-producing states have significant effects on manufactured exports but not on agricultural exports. Additionally, Iran-Iraq war does not have significant effect on the exports from either sector, whereas the rainfall variable is a significant predictor of agricultural exports. They concluded from their findings that the Turkish boom was mostly the result of adopting appropriate policy measures but external circumstances gave also additional boost to exports. Another study by Uygur (1997) makes a comprehensive evaluation of export policies pursued in Turkey during the period from the late 1970s to mid-1990s. In this study an export supply equation is constructed to assess the effects of exchange rate policies, different export promotion schemes and domestic demand policies. Another focus is to uncover the differences of these policies in terms of their long-term and short-term effects. The findings obtained from dynamic error correction estimation of an export supply equation show that real exchange rate is the most significant variable both in the short-term and long-term. Domestic demand has also significant effect on export supply. The effect of export subsidies on export supply is significant and positive in the short-term but it turns out to be negative in the long-term.

Şahinbeyoğlu and Ulaşan (1999) estimated export supply and demand functions for Turkey covering the period 1987-1998. Their primary aim was to analyze the validity of historical studies as a reliable guide to the future trends in export. They considered real export supply as a function of real domestic income and real effective exchange rate. Real export demand is considered as a function of real foreign income and real effective exchange rate. According to the error correction model (ECM) estimation results, both price and income elasticities of real export demand and supply functions are less than one, referring to inelastic components. In addition to ECM estimation results, Chow breakpoint test reveals that both long-run and short-run elasticities are stable during the sample period. Lall (2000) considered another aspect of Turkish export. The author investigated the position and prospects of Turkish manufactured exports by analyzing its technological structure. He concluded that the structure of exports is extremely weak and exports comprise essentially low technology products. He investigated also the feasibility of several high-technology growth strategies for Turkey. Özatay (2000) constructed a quarterly macro econometric model representing the main features of the Turkish economy. In this respect, he estimated total exports of goods as a function of foreign income, real exchange rate and long-run relation between real exports and real exchange rate. Estimation results show that foreign income variable does not have a significant coefficient. Real exchange rate is statistically significant and has a coefficient less than one.

Senhadji and Montenegro (1999) estimated export demand elasticities for a large number of developing and industrial countries including Turkey. Real exports of the home country are determined as a function of the export price of the home country relative to the price of its competitors and the activity variable defined as real GDP minus real exports of the home country's trading partners. According to their results, for Turkey, the short-run price and foreign income elasticities of demand are less than one, whereas the long-run price elasticity of demand is far greater than one. The long-run foreign income elasticity of demand is still less than one. Atabek and Çevik (2001) implemented a comprehensive investigation of Turkey's trade performance. They estimated import demand function in addition to export supply and demand functions. One of the main conclusions of the study is that both export supply and export demand is price (exchange rate) inelastic but income elastic in the long-run whereas it is price insensitive and income inelastic in the short-run. One of the recent studies investigating Turkey's trade flows is the study of Vehbi (2002). The author constructed an error-correction model (ECM) to forecast exports and imports of Turkey in the broad economic categories (BEC) classification. The main findings of the study indicate that economical conjunctures of main trading partners are more significant than the price changes on export performance.

Terzi and Zengin (1999) emphasized the importance of investigating the relationship between real exchange rate, imports and exports on sectoral disaggregation. They claimed that estimation of disaggregate data might reveal some relationships that cannot be observed in aggregate level. Following a similar view, in this study, exports are disaggregated on the basis of main trading partners of Turkey and major industrial classification of commodities. Investigation of export performance on sectoral basis is essential to monitor the structure of the export. Total export and disaggregated export based on the data of major trading partners are estimated for the period 1989-2000. Sectoral export demand functions are estimated for the period 1994-2000. Aggregate export demand equation is estimated using fixed effect models whereas sectoral and country specific export demand equations are estimated using seemingly unrelated regression (SUR) estimation methodology. Each equation is estimated using panel data techniques, so that variation over both the cross section and time series dimensions are jointly considered.

The plan of the paper is as follows: In Section 2, a brief review of Turkish export is presented. Section 3 gives information on the data and the model used. Section 4 discusses the panel data techniques. In the following section, an application for Turkey is given whereas in Section 6 export demand elasticities based on selected sectors and countries are presented. Finally, Section 7 summarizes main conclusions of the study.

2. Export Performance in Turkey

Turkey followed an inward oriented development strategy up to 1980s and faced several external and internal shocks mainly due to high increases in oil prices. As a result of these crises, trade and current account deficits increased sharply, economic growth slowed down and inflation rate raised. Accordingly, a stabilization program was introduced in 1980 with the aim of reducing external deficit and inflation rate. The main impact of the implemented program was the reduction in domestic demand, which led to an increase in exports, among other effects (like export credit expansion and large devaluations).

During the 1981-1988 period, exports increased and the Turkish economy experienced an export-led growth. But due to the expansionary monetary policies and the appreciation of the Turkish lira, export performance slowed significantly during the period 1989-1993. Another 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 Southeast Asia and Russian Federation. The earthquakes occurred in 1999 also affected the economic conditions negatively.

In analyzing the export performance, the structure of exports has to be analyzed. Table 1 presents percentages of Turkey's main trading partners in total exports. As it can be seen from table, Germany has the highest share in our total exports. USA and UK follow Germany.

Table 1
Shares of main trading partners in total exports (%)

	1996	1997	1998	1999	2000
USA	7.1	7.7	8.3	9.2	11.3
Germany	22.3	20.0	20.2	20.6	18.6
France	4.5	4.4	4.8	5.9	6.0
Netherlands	3.3	3.0	3.3	3.5	3.1
UK	5.4	5.8	6.4	6.9	7.3
Italy	6.2	5.3	5.8	6.3	6.4
Total exports	100.0	100.0	100.0	100.0	100.0

Source: Undersecretariat of Foreign Trade.

As a result of the export-led growth strategies implemented in 1980s, the structure of exports has changed from primary products to more technology-intensive products. The share of industrial products in total exports increased whereas the share of primary products decreased in these years. Table 2 gives share of main commodity groups in total exports for the years 1996-2000. As it can be seen from Table 2, export of manufactures is concentraded on several products like textiles, clothing, machinery and transport equipment. Iron and steel has also a significant proportion in total exports. Bearing in mind that, iron and steel, textiles, clothing and food sectors are classified mostly as semi-processed primary goods, it can be said that Turkish exports is mainly dependend on low-technology products. But in order to obtain a sustainable export growth, the structure of exports has to be changed in favor of technology-intensive products. Textiles and clothing sectors together account for nearly 40 percent of total exports. This shows that diversification of exports has not been achieved yet.

Table 2
Share of disaggregated exports in total exports (SITC-Rev.3, %)

	1996	1997	1998	1999	2000
1- AGRICULTURAL PRODUCTS	21.3	20.8	18.7	16.7	13.9
i-Food	19.6	19.5	17.4	15.4	12.8
ii-Agricultural Raw Materials	1.7	1.3	1.4	1.3	1.1
2- MINING PRODUCTS	4.3	3.8	3.8	4.1	4.2
i- Metalliferous ores and metal scrap	1.8	1.8	1.5	1.6	1.6
ii- Mineral fuels, lubricants and related materials	1.2	0.7	1.0	1.3	1.2
iii-Non-ferrous metals	1.3	1.3	1.4	1.2	1.3
3- MANUFACTURES	74.3	75.3	77.4	79.1	81.7
i-Iron and steel	8.3	8.6	6.8	6.5	6.7
ii-Chemicals	4.3	4.5	4.3	4.2	4.5
iii-Other semi-manufactures	6.9	7.1	7.5	7.7	8.2
iv- Machinery and transport equipment	13.0	12.8	15.2	18.9	20.7
v- Textiles	11.7	12.8	13.2	13.1	13.3
vi- Clothing	26.2	25.5	26.2	24.5	23.7
vii - Other consumer goods	4.0	4.1	4.2	4.1	4.6
4- OTHER PRODUCTS	0.1	0.1	0.1	0.2	0.2
TOTAL	100.0	100.0	100.0	100.0	100.0

Source: Undersecretariat of Foreign Trade

3. Data and the Model

In this study, the imperfect substitutes model proposed by Goldstein and Khan (1985) is followed. The major assumption of this model is that neither imports nor exports are perfect substitutes for domestic goods. Exports are imperfect substitutes in world markets for other countries' domestically produced goods, or for third countries' exports. The conventional demand theory says that, the consumer is postulated to maximize utility subject to a budget constraint. In this respect, export demand function is specified as a function of the real exchange rate and the rest-of-world real incomes. Thus, the export demand equation can be expressed as:

$$\log E_{i,t} = \alpha_i + \beta_1 \log I_{i,t} + \beta_2 \log R_{i,t} + u_{i,t}, \quad i=1,\ldots,6\,; \quad t=1,\ldots,47 \qquad (3.1)$$

In the equation given above, i denotes the six countries which are the most important trade partners of Turkey (Germany, USA, Italy, UK, France and Netherlands) and t denotes the time. The description of the variables are listed below:

- $E_{i,t}$: export of Turkey to country i;

- $I_{i,t}$: volume index of Gross Domestic Product (GDP) of country i (a proxy for the foreign income);

- $R_{i,t}$: real exchange rate of country i.

The formula of real exchange rate is given as:

$R_{i,t} = \dfrac{P_d}{e_i * P_i}$ where,

- e_i : nominal exchange rate of country i;

- P_d : price of domestic goods (Wholesale Price Index)

- P_i : producer prices of country i.

In equation (3.1), β_1 is the real foreign income elasticity and β_2 is the real exchange rate elasticity of export demand. Based on the theory, it is expected that, β_1 has a positive sign, indicating that demand rises as income increases and β_2 has a negative sign, implying an increase in demand with the depreciation of Turkish Lira (TL). The dependent variable is deflated by export price index. The model estimations are based on quarterly data between the years 1989-2000, the base year being 1987. The data is obtained from the database of IMF and Central Bank of Turkey. The export demand equation is estimated using panel data techniques.

The benefits from using panel data estimation are various. In panel data estimation, variations over both the cross-section and time series dimensions are considered jointly. This brings the advantage of using all the information available which are not detectable in pure cross-sections or in pure time series data. In addition to this, panel data estimation provides improved coefficient estimates by increasing the power of the tests. Because of these and many more advantages, panel data techniques are employed in this study.

In panel regression, different models, like one-way and two-way error correction models, can be constructed according to the structure of the error term. In one-way error component regression model, there is only one effect, which can be either individual effects or time effects. But in two-way error component regression model there are both effects. In one-way error component model, μ_i denotes the unobservable individual specific effects whereas v_{it} denotes the remainder disturbances ($u_{it} = \mu_i + v_{it}$). On the other hand, in the two-way error component regression model, μ_i denotes the unobservable individual specific effects, λ_t denotes the unobservable time effect and v_{it} denotes the remainder stochastic disturbance term. In order to determine which model is appropriate, the existence of individual and/or time effects can be tested. In this case, the null hypotheses tested are:

$$H_{01}: \sigma_\mu^2 = \sigma_\lambda^2 = 0 \quad \text{(no time and individual effects)} \tag{3.2}$$

$$H_{02}: \sigma_\mu^2 = 0 \quad \text{(no individual effects)} \tag{3.3}$$

The hypotheses are tested using F-test. Estimation of variance components may have different forms in accordance with the assumption made about the error component model (whether there is individual effects or time effects). In error components model, different models that change according to the covariance structure of the residuals can be used. The fixed effects and random effects models are two of them. In the fixed effects models, the parameter μ_i is assumed to be fixed. The remainder stochastic disturbance term, v_{it}, is independently, identically distributed. In order to the test the significance of the fixed effects model, the dummy variables representing the individual effects are tested jointly.[1]

[1] Using standard Wald-F statistic test, the regression equation given above is tested. In the random effects model, the parameters μ_i and v_{it} are assumed to be independently, identically distributed. The statistical significance of the random effects are tested by using the Breusch-Pagan LM test statistic. In order to decide if there is a statistical difference between the fixed and random effects, the Hausman specification test can be applied. The Hausman test indicates that the difference between the fixed and random effects models is in their respective covariance matrices. The Hausman test statistic is distributed as a Chi-square with degrees of freedom equal to the number of slope parameters.

The identification of fixed and random effects models has been widely discussed in the literature. When N individuals are randomly drawn from a large population, random effects model is appropriate. Whereas, when the interest is on specific N individuals, fixed effects model is appropriate. The countries under investigation are the main purchasers of the Turkish exported goods. Since, they can represent the structure of the Turkish export very well. So, the choice of the countries is not random. Although, the result of the Hausman test indicates no significant difference between fixed effects and random effects, fixed effects model is used in this study due to the reasons given above[2].

4. The Econometric Methodology

The stationarity of variables in the model (3.1) is tested using the ADF unit root test procedure. After that, panel unit root tests are applied. In recent years some tests for unit roots and cointegration within panels are developed in the literature. Levin and Lin (1992, 1993), Im, Pesaran and Shin (1997), Maddala and Wu (1999), Kao (1999) and Quah (1994) have developed unit root tests within panels. The panel cointegration tests developed by Kao and Chiang (1998), Pedroni (1995, 1997), McCoskey and Kao (1998), Phillips and Moon (1999) are also widely used. In this study panel unit root tests of Levin and Lin (hereafter LL) and Im, Pesaran and Shin (hereafter IPS) are used. LL test is preferred because of its large potential power gains. Besides, LL test is widely used in empirical researches. IPS test is the extension of LL unit root test.

Panel unit root tests: Conventional unit root tests examine the unit-root null based on a single equation method. LL (1992) showed that, applying a unit root test on a pooled cross-section data set, rather than performing separate unit-root tests for each individual series, can increase statistical power. Wu (1996), Oh (1996),

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W= $(b_{random} - b_{fixed})'[Var(b_{random}) - Var(b_{fixed})]^{-1}(b_{random} - b_{fixed})$

In the equation given above, b_{random} demonstrates the estimated slope parameters of the random effects model, b_{fixed} demonstrates the estimated slope parameters of the fixed effects model and $Var(b_{random})$ and $Var(b_{fixed})$ demonstrates the estimated covariance matrices for the random and fixed effects models respectively.

[2] Matyas, L., Sevestre, P., (1996): 'The Econometrics of Panel Data', *Kluwer Academic Publishers.*

MacDonald (1996) and Frankel and Rose (1996) are some of the examples who argue that the commonly used unit root tests like Dickey-Fuller (DF), augmented Dickey-Fuller (ADF) and Phillips-Perron (PP) tests lack power in distinguishing the unit root null from stationary alternatives. They also claimed that using panel data unit root tests is one way of increasing the power of unit root tests based on a single time series.

LL (1992) panel unit root test is based on the following regression:

$$\Delta y_{i,t} = \rho y_{i,t-1} + \alpha_0 + \delta t + \alpha_i + \theta_t + \varepsilon_{i,t} \ , \quad i=1,2,\ldots,N \ , \quad t=1,2,\ldots,T \tag{4.1}$$

The model incorporates a time trend as well as individual and time specific effects. Levin and Lin considered six subcases of model 4.1. All the models are estimated using OLS as a pooled regression model. The differences of the submodels lie in the specification of the regression equation (i.e. the inclusion of individual specific intercepts and time trends). The first one includes neither intercept nor time trend whereas the last one includes both intercept and time trend vary with individuals.

LL (1993) provided some new results on panel unit root tests. LL unit root test is improved to deal with the problem of heteroscedasticity and autocorrelation. The panel test procedure involves several steps.

- First, in order to remove the influence of aggregate effects, the cross-section averages are subtracted from each variable.

$$\overline{y}_t = \frac{1}{N}\sum_{i=1}^{N} y_{i,t}$$

- Second, ADF test is applied to each individual series and the disturbances are normalized. For example, the regression equation under investigation, i.e. the model given in (4.2) is estimated for each individual with $\Delta y_{i,t}$ and $y_{i,t-1}$ being dependent variable respectively.

$$\Delta y_{i,t} = \alpha_i + \rho_i y_{i,t-1} + \gamma_i t + \sum_{j=1}^{p_i} \theta_{i,j} \Delta y_{i,t-j} + \varepsilon_{i,t} \ , \quad t=1,2,\ldots,T \tag{4.2}$$

The residual terms obtained from these two auxiliary regressions are denoted by $\hat{e}_{i,t}$ and $\hat{V}_{i,t-1}$ respectively. Then $\hat{e}_{it}$ are regressed on $\hat{V}_{i,t-1}$:

$$\hat{e}_{i,t} = \rho_i \hat{V}_{i,t-1} + \varepsilon_{i,t} \tag{4.3}$$

to get $\hat{\rho}_i$ which is equivalent to the OLS estimator of ρ_i in 4.2 directly. To eliminate the heteroscedasticity in $\varepsilon_{i,t}$, Levin and Lin suggest the following normalization to control for heterogeneity across individulas:

$$\hat{\sigma}_{e_i}^2 = \frac{1}{T - p_i - 1} \sum_{t=p_i+2}^{T} (\hat{e}_{i,t} - \hat{\rho}_i \hat{V}_{i,t-1})^2 \tag{4.4}$$

$$\tilde{e}_{i,t} = \hat{e}_{i,t} / \hat{\sigma}_{e_i} \tag{4.5}$$

$$\tilde{V}_{i,t-1} = \hat{V}_{i,t-1} / \hat{\sigma}_{e_i} . \tag{4.6}$$

- Third, the ratio of long-run to short-run standard deviation for each individual series are estimated and the average ratio for the panel is calculated as:

$$\hat{S}_{NT} = \frac{1}{N} \sum_{i=1}^{N} \frac{\hat{\sigma}_{y_i}}{\hat{\sigma}_{e_i}} \tag{4.7}$$

where the long-run variance of $\hat{\sigma}_{y_i}^2$ is estimated by

$$\hat{\sigma}_{y_i}^2 = (T-1)^{-1} \sum_{t=2}^{T} \Delta y_{i,t}^2 + 2 \sum_{L=1}^{\bar{K}} w_{\bar{K}L} \left(\frac{1}{T-1} \sum_{t=L+2}^{T} \Delta y_{i,t} \Delta y_{i,t-L} \right) \tag{4.8}$$

$\bar{K}$ is the lag truncation parameter and $w_{\bar{K}L}$ is some lag window. The estimated average standard deviation ratio $\hat{S}$ will be used to adjust the mean of the panel unit root test statistic.

- The next step is to compute the panel test statistic. Under the null hypothesis, the normalized residual innovations $\tilde{e}_{i,t}$ are independent of the normalized lagged residuals $\tilde{v}_{i,t-1}$ for each individual in the panel. This hypothesis can be tested by performing the following regression:

$$\tilde{e}_{i,t} = \rho \tilde{V}_{i,t-1} + \tilde{\varepsilon}_{i,t} \tag{4.9}$$

using a total of $N\tilde{T}$ observations. For the regression equation given above, the t-statistics is given by:

$$t_{\rho=0} = \frac{\hat{\rho}}{RSE(\hat{\rho})} \tag{4.10}$$

where

$$RSE(\hat{\rho}) = \hat{\sigma}_{\varepsilon} \left[\sum_{i=1}^{N} \sum_{t=p_i+2}^{T} \tilde{V}_{i,t-1}^2 \right]^{-1/2} \tag{4.11}$$

$$\hat{\sigma}_{\varepsilon}^{2} = (N\tilde{T})^{-1} \sum_{i=1}^{N} \sum_{t=p_i+2}^{T} (\tilde{e}_{i,t} - \hat{\rho}\tilde{V}_{i,t-1})^2 \tag{4.12}$$

$$\widetilde{T} = (T - \overline{p} - 1) \text{ and } \overline{p} = N^{-1} \sum_{i=1}^{N} p_i$$

$\overline{p}$ is the average lag length used in the individual ADF regression. Since the test statistic is not centered at zero, Levin and Lin suggest using the following adjusted t-statistic:

$$t_{\rho}^{*} = \frac{t_{\rho=0} - N\tilde{T}\hat{S}_{NT}\hat{\sigma}_{\varepsilon}^{-2} RSE(\hat{\rho})\mu_{\tilde{T}}^{*}}{\sigma_{\tilde{T}}^{*}}. \tag{4.13}$$

The mean and standard deviation adjustment terms ($\mu_{\tilde{T}}^{*}$ and $\sigma_{\tilde{T}}^{*}$) are obtained using Monte Carlo simulation technique and their values are given in LL (1993).

Under the null hypothesis;

$$H_0 : \rho = 0,$$

the adjusted test statistic t_{ρ}^{*} has a $N(0,1)$ distribution and the critical values of the standard normal distribution can be used to test the null hypothesis that

$$\rho_i = 0 \text{ for all i=1,…,N.}$$

In this paper, LL panel unit root test is performed by following the methodology given in Rapach (2001). According to this methodology, the author performs a Monte Carlo simulation to obtain the critical values when testing the unit root null hypothesis that $\rho = 0$. The methodology proceeds as follows: First, the regression given in (4.2) is estimated by using OLS under the unit root null hypothesis by restricting ρ and γ_i to be equal to zero for the panel. Then using the restricted OLS estimates of α_i and $\theta_{i,j}$, random draws from a $N(0,\hat{\sigma}^2)$, where $\hat{\sigma}^2$ is the restricted OLS estimate of σ^2, and setting the initial $y_{i,t-1}$ and $\Delta y_{i,t-j}$ values equal to zero, a panel series of T+100 observations for $\Delta y_{i,t}$ is simulated. The first 100 observations control for initial value bias and are discarded to form a series of T observations, corresponding to the original sample. The process is repeated 2000 times to generate 2000 simulated panel series. For each panel series, the t_{ρ} statistic is calculated and stored. The t_{ρ} statistics are then ordered and the 20th, 100th and 200th values serve as the 1, 5 and 10 percent critical values respectively.[3]

[3] Rapach D. E. (2001). The program to perform LL test written in GAUSS code is obtained from the author's home page.

Different from the LL test, IPS panel unit root test contains heterogeneous adjustment processes and pool the t-statistics from univariate independent ADF regressions. IPS relax the assumption that $\rho_1 = \rho_2 = \ldots = \rho_N$ under H_1. Like the LL test, cross-sectional means are subtracted from the data to remove the common time specific effects. The ADF regression equation given in (4.2) is estimated for N individual separately and then the average of the t-statistics are calculated. Let $t_{i,T}$ denote the t-statistic for testing unit roots, then the test statistic is denoted as:

$$\sqrt{N}\frac{\left(\bar{t}_{N,T} - \mu\right)}{\sigma} \Rightarrow N(0,1), \text{ where } \bar{t}_{N,T} = \frac{1}{N}\sum_{i=1}^{N} t_{i,T} \quad . \tag{4.14}$$

The values of μ and σ are calculated by using Monte Carlo methods and are given in IPS (1997). One limitation of the IPS test is related with the lag length of the dependent variable in the ADF regression. In the case of serial correlation, IPS propose using the ADF t-test with different lag length for individual series. However, μ and σ will vary as the lag length included in the ADF regression varies. And, to make use of their tables of μ and σ, one is restricted implicitly to using the same lag length for all the ADF regressions for individual series. The comparison of LL and IPS panel unit root tests yields some broad comments on the merits and demerits of these methods. Some of these can be listed as follows:

-The LL test tests a very restrictive hypothesis that is rarely of practical interest.

-One drawback to the test is that the first order autoregressive coefficient (ρ_i) is restricted to be identical across countries under the null and alternative hypothesis. But IPS allow ρ to differ across countries under the alternative hypothesis.

-A potential problem with both the LL and IPS panel tests is the cross-sectional dependence. In order to overcome the cross-sectional dependence, LL and IPS propose subtracting the cross-sectional means from both sides of (4.2) prior to estimation. However, O'Connell (1998) shows that this procedure will do little to reduce cross-sectional dependence.

-In the IPS test, the distribution of the t-bar statistic involves the mean and variance of the t-statistics used. IPS compute the values of mean and variance for the ADF test statistic for different values of the number of lags used and different sample sizes. However, these tables are valid only if the ADF test is used for the unit root tests. Also, if the length of the time series for the different samples is different, there is a problem using the tables prepared by IPS (Maddala, Wu (1999)).

-IPS test has better small sample properties than the LL test and has the additional advantage of simplicity.

Panel cointegration test: There are different methods for testing cointegration in panels. The first method takes the null hypothesis of no cointegration and uses residuals derived from the panel regression of Engle and Granger (1987) method. Pedroni (1995, 1997), McCoskey and Kao (1998) panel cointegration tests are based on this method. Another approach is to take the null of cointegration and is the basis of the tests proposed by Harris and Inder (1994), Shin (1994), Leybourne and MacCabe (1994) and Kwiatowski *et. al*. (1992). All the panel data cointegration tests allow for heterogeneity in the cointegrating coefficients. But one drawback related with these tests is that the null and alternative hypotheses imply that either all the relationships are cointegrated or all the relationships are not cointegrated. Except the Fisher (1932) test, there is no allowance for some relationships to be cointegrated and others not. Initially developed panel cointegration tests applied panel unit root tests directly to the residuals from an Engle Granger type two-step methodology. But the recent opinion in the literature suggested that the test statistics using this approach would be biased towards accepting stationarity. Pedroni (1995) shows that applying panel unit root tests directly to regression residuals is inappropriate for several reasons like the lack of exogeneity of the regressors and the dependency of the residuals on the distribution of the estimated coefficients (see Pedroni (1995,1997) for details). For these reasons it is important to have a test procedure for cointegration which is robust to the presence of heterogeneity in the alternative. Since the cointegration test proposed by Pedroni allows for considerable heterogeneity, it is preferred in this study. The cointegrating system considered is given as follows:

$$y_{i,t} = \alpha_i + \delta_i t + \gamma t + \beta_{1i} x_{1i,t} + \beta_{2i} x_{2i,t} + \ldots + \beta_{Mi} x_{Mi,t} + e_{i,t}\,, \quad (4.15)$$

$$t=1,\ldots,T;\ i=1,\ldots,N;\ m=1,\ldots,M$$

where T is the number of observations over time, N is the total number of individual units in the panel and M is the number of regression variables. In the equation (4.15), α_i is the member specific intercept, γ_t is a time dummy common to members of the panel and $\delta_i t$ is the deterministic time trend, which are specific to individual panel members.

Pedroni discusses the construction of seven panel cointegration statistics, four based on pooling along the within-dimension and three based on pooling along the between-dimension. Within the first category, three of the four tests involve the use of non-parametric corrections familiar from the work of Phillips and Perron (1988). The fourth is a parametric ADF-based test. In the second category, two of the three tests use non-parametric corrections while the third is again an ADF-based test. The test statistics given in the first category are based on estimators that effectively pool the autoregressive coefficient across different members for the unit root tests on the estimated residuals, while the test statistics given in the second category are based on estimators that simply average the individually estimated coefficients for each member i.[4]

The existence of the cointegration relationship between the variables is investigated through the stationarity of the error term in equation (4.15). For the non-parametric tests, the constructed equation is given as:

$$\hat{e}_{i,t} = \rho_i \hat{e}_{i,t-1} + \hat{u}_{i,t} \tag{4.16}$$

whereas the parametric tests estimate:

$$\hat{e}_{i,t} = \rho_i \hat{e}_{i,t-1} + \sum_{k=1}^{k_i} \hat{\rho}_{i,k} \Delta \hat{e}_{i,t-k} + \hat{u}_{i,t} \, . \tag{4.17}$$

For the first category, the null of no cointegration is given as:

$$H_0 : \rho_i = 1 \text{ for all individuals}$$

$$H_1 : \rho_i = \rho < 1 \text{ for all individuals}$$

For the second category, the null of no cointegration is given as:

$$H_0 : \rho_i = 1 \text{ for all individuals}$$

$$H_1 : \rho_i < 1 \text{ for all individuals}$$

The alternative hypothesis given in the second category does not presume a common first order autoregressive coefficient for all individuals. The test statistics in both categories have asymptotically standard normal distribution;

$$\frac{x_{N,T} - \mu\sqrt{N}}{\sqrt{v}} \Rightarrow \text{N (0,1)} \tag{4.18}$$

[4] P. Pedroni, Critical Values for Cointegration Tests, Oxford Bulletin of Economics and Statistics (1999)

where $x_{N,T}$ is the form of the tests statistic. The values of μ and ν are the corresponding values for each test of the mean and variance respectively and given in Pedroni (1999). There is one issue which is important to keep in mind when performing Pedroni panel cointegration tests. It is known that the cointegrating vectors are not unique in general. However, Pedroni does not address the issue of normalization, how to establish the number of cointegrating relationships or how many cointegrating relationships exist among a certain set of variables. The assumption of these tests is that the researcher has in mind a particular normalization among the variables which is deemed sensible and is simply interested in knowing whether or not the variables are cointegrated (Pedroni (1999)).

5. Empirical Results

The null hypothesis (3.2) and (3.3) given in Section 2 are tested using F-statistic. At the 5 percent significance level, the null hypothesis H_{01}, indicating the joint significance of time and individual effects, is accepted. It means time and individual effects are not jointly significant. At the same time, the null hypothesis H_{02} is rejected at the 5 percent significance level, indicating the significance of individual effects. In this respect, applying the fixed effects model in one-way error component regression model found to be appropriate.

As a first step, the stationarity of each series is investigated by applying ADF unit root test. After that LL and IPS panel unit root tests are performed.

Table 3
ADF Unit Root Test

	ADF test stat.	K	1% critical value	5% critical value
E_USA	-2.750	3,T	-4.184	-3.516
I_USA	-1.785	1,T	-4.173	-3.511
R_USA	-3.288	3,T	-4.184	-3.516
E_Germany	-2.342	1,C	-3.581	-2.927
I_Germany	-2.394	1,C	-3.581	-2.927
R_Germany	-2.032	1,C	-3.617	-2.942
E_France	-2.318	2,T	-4.178	-3.514
I_France	-2.249	4,C	-3.593	-2.932
R_France	-2.027	2,C	-3.585	-2.929

	ADF test stat.	K	1% critical value	5% critical value
E_Italy	-2.778	1,T	-4.173	-3.511
I_Italy	-1.933	1,T	-4.173	-3.511
R_Italy	-4.134*	2,T	-4.178	-3.514
E_Netherlands	-2.377	1,C	-3.581	-2.927
I_Netherlands	-2.338	2,C	-3.585	-2.929
R_Netherlands	-3.162	3,T	-4.184	-3.516
E_UK	-3.610*	1,T	-4.173	-3.511
I_UK	-2.616	1,T	-4.173	-3.511
R_UK	-3.870*	3,T	-4.184	-3.516

-K denotes the highest order of lag for which t-statistic in the regression is significant. C denotes a significant intercept and T denotes a significant intercept and time trend.
-The lag lengths are chosen according to Akaike Information Criteria (AIC).
-(*) denotes a significant coefficient at 5 % critical value.

The regression equations given in Table 3 involve intercept and time trend. According to the MacKinnon critical values, for almost all series, the null of unit root cannot be rejected at 1 and 5 percent significance level.

Table 4
Panel unit root test

Series	k	LL	LL panel ADF test critical values 1%	5%	10%	IPS
Export	0	-13.18*	-6.77	-6.14	-5.86	-1.56
	1	-11.71*	-6.70	-6.13	-5.85	
	2	-7.08*	-6.61	-6.02	-5.71	
	3	-4.53	-6.54	-5.95	-5.64	
Income	0	-4.81	-6.77	-6.12	-5.83	-0.12
	1	-4.92	-6.53	-5.91	-5.60	
	2	-5.21	-6.33	-5.75	-5.49	
	3	-5.67**	-6.46	-5.88	-5.60	
Real exc. Rate	0	-5.85	-6.72	-6.19	-5.88	0.44
	1	-6.13**	-6.85	-6.17	-5.89	
	2	-5.93**	-6.66	-6.07	-5.75	
	3	-5.22	-6.60	-6.06	-5.76	

-k is the order of the lagged dependent variable in the equation (4.2).
-IPS unit root test statistic is calculated by using mean and variance values given in Im, Pesaran and Shin (1997).
-The lag lengths are chosen according to Akaike Information Criteria (AIC).
-(*) denotes significance at 1 % critical value
-(**) denotes significance at 10 % critical value

The panel unit root test results reported in Table 4 shows evidence of nonstationarity. According to LL test results, the null of nonstationarity is accepted for all the critical values at lag order 3. For real income and real exchange rate series the null of nonstationarity is accepted at one and five percent critical values at all lag orders.

The univariate ADF unit root test and panel data unit root tests are also performed for first differenced series. Although the results are not reported here, real export, real foreign GDP and real exchange rate for all individuals are found to be integrated of order 1. One of the ways to deal with I(1) variables is to investigate the cointegration relationship between variables. The Johansen cointegration test results for the six countries are given at Table 5. The third and fourth columns of Table 5 give the eigenvalues and likelihood ratio statistics. The second column presents the results of the null hypothesis of no cointegration and at most one cointegration relationship between variables. The hypothesis that there is at most one cointegration relationship between the variables is accepted for all countries except USA and Germany. For these countries two cointegrating vectors are found.

Table 5
Johansen cointegration test

	Null hypothesis	Eigenvalue	Likelihood ratio	5 Percent Critical Value	1 Percent Critical Value	Num. of coint. equations
USA	Ho: r=0 **	0.63	63.28	29.68	35.65	2
	Ho: r<=1*	0.26	18.00	15.41	20.04	
Germany	Ho: r=0 **	0.61	55.77	34.91	41.07	2
	Ho: r<=1*	0.34	20.62	19.96	24.60	
France	Ho: r=0 *	0.38	29.95	29.68	35.65	1
	Ho: r<=1	0.14	8.79	15.41	20.04	
Netherlands	Ho: r=0 **	0.49	44.65	29.68	35.65	1
	Ho: r<=1	0.24	14.21	15.41	20.04	
Italy	Ho: r=0 *	0.39	30.33	29.68	35.65	1
	Ho: r<=1	0.17	8.29	15.41	20.04	
UK	Ho: r=0 **	0.42	42.35	34.91	41.07	1
	Ho: r<=1	0.25	18.21	19.96	24.60	

- (*) denotes rejection of null hypothesis at 5% significance level
- (**) denotes rejection of null hypothesis at 1% significance level
- r indicates the number of cointegrating vectors
- For the countries except Germany and UK an assumption of linear deterministic trend in data is allowed. For all countries seasonal dummies are added as exogenous variables.

Table 6
Pedroni panel cointegration test

Panel v-statistic	Panel rho-statistic	Panel pp-statistic	Panel adf-statistic	Group rho-statistic	Group pp-statistic	Group adf-statistic
0.786	-4.754	-5.884	-1.128	-6.959	-9.503	-3.087

Except panel variance and panel ADF statistics, all of the panel cointegration test statistics developed by Pedroni rejects the null of no cointegration at 5 percentage significance level. Since there is a cointegration relationship between the variables, the Engle and Granger two-step method can be used. According to Engle and Granger (1987), if the variables are cointegrated, the stable long-run relationship can be estimated by standard least-squares techniques. The Engle and Granger method consists of two steps. In the first step, the regression equation given below is estimated to obtain the long-run coefficients α_i, β_1 and β_2.

$$\log E_{i,t} = \alpha_i + \beta_1 \log I_{i,t} + \beta_2 \log R_{i,t} + e_{i,t} \tag{5.1}$$

In the second step, stationarity of the residuals of the estimated equations are tested by the ADF test. According to the test results given in Table 7, residuals of all equations are stationary.

Table 7
Engle-Granger cointegration test results

	ADF test result	Lag order
RESID_USA	-5.4183**	0
RESID_Germany	-4.9952**	0
RESID_France	-4.7585*	0
RESID_Netherlands	-5.7411**	0
RESID_Italy	-4.9087**	0
RESID_UK	-4.9030**	0

- (*) denotes rejection of null hypothesis of unit root at 5% significance level.
- (**) denotes rejection of null hypothesis of unit root at 1% significance level.
- The critical values for this test are taken from Engle and Yoo (1987) for 50 observations and 3 variables.

Engle-Granger cointegration test and Johansen cointegration test give evidence in favor of cointegration relationship between variables for all countries. Based on these results, following error-correction model is estimated.

$$\Delta \log E_{i,t} = \delta_i + \beta_1 \Delta \log I_{i,t} + \beta_2 \Delta \log R_{i,t} + \lambda \left[\log E_{i,t-1} - \hat{\alpha}_i - \hat{\beta}_1 \log I_{i,t-1} - \hat{\beta}_2 \log R_{i,t-1} \right] + e_{i,t} \quad (5.2)$$

The coefficient of the error-correction term, λ, represents the speed of adjustment to the long-run relationship estimated in the first step. The model can be estimated using the fixed-effects model. In order to allow for lagged adjustment, lagged dependent variables of regressors both in long-run and short-run equations are allowed. The coefficients α_i and δ_i (in equations (5.1) and (5.2)) are different for each individual whereas β_1 and β_2 are the same for all individuals due to the fixed effects model. The long-run and short-run estimation results are given at Table 8 and Table 9 respectively. In both long-run and short-run, the signs of the coefficients of foreign GDP are as expected. But in the long-run, the coefficient of the real exchange rate at time t is found to be positive. This indicates a decrease in the export due to the depreciation of Turkish lira in the long-run. Bearing in mind that Turkish manufacturing industry is highly dependent on imports of intermediate goods, positive coefficient of real exchange rate will be more clear. Depreciation of Turkish lira may cause a decrease in the imports and as a result of this a decrease in the production and exports.

The long-run elasticity of export demand with respect to the foreign GDP and real exchange rate are estimated to be 4.53 and 0.42 respectively. Whereas the short-run elasticity of export demand with respect to the foreign GDP and real exchange rate are estimated to be 3.84 and -0.58 respectively. The long-run elasticity of export demand with respect to the real exchange rate is lower than the short-run elasticity. Whereas the short-run elasticity of export demand with respect to the foreign GDP is lower than the long-run elasticity. The foreign income elasticity is found to be near 4 for real export demand. However, real exchange rate elasticity is lower than one referring to inelastic component. According to this result, it can be argued that Turkish exports can be mainly explained by foreign income (or foreign demand) changes rather than real exchange rate changes from the demand side. The results differ from Atabek and Çevik (2001) and Şahinbeyoğlu and Ulaşan (1999) in the respect of income elasticity. Atabek and Çevik found that export demand is income elastic in the long-run but income inelastic in the short-run. Şahinbeyoğlu and Ulaşan found the income elasticity of export demand less than one both in the long-run and short-run. In this study export demand is found to be price insensitive. These results give support to the hypothesis that exchange rate policies may not be successful in promoting export growth.

Table 8
Estimation of the long-run model

Variable	*Coefficient*	*t-Statistic*	*p-value*
$\log I$	4.531	26.553	0.000
$\log R$	0.425	2.458	0.015
$\log R(-1)$	-0.475	-2.705	0.007
Fixed Effects			
USA—C	-7.1064		
Germany--C	-0.0796		
France—C	-7.1887		
Netherlands--C	-7.9103		
Italy—C	-6.8316		
UK—C	-7.1465		

-Estimation is corrected for both cross-section heteroskedasticity and contemporaneous correlation.

Table 9: Estimation of the short-run model

Variable	*Coefficient*	*t-Statistic*	*p-value*
$\Delta \log I$	3.843	3.381	0.001
$\Delta \log R(-1)$	-0.582	-3.412	0.001
ECM(-1)	-0.611	-10.519	0.000
Fixed Effects			
USA--C	0.0062		
Germany--C	0.0072		
France--C	0.0078		
Netherlands--C	-0.0187		
Italy--C	0.0030		
UK--C	0.0080		

-Estimation is corrected for both cross-section heteroskedasticity and contemporaneous correlation

- Δ denotes first difference.

6. Income, Price and Production Elasticities of Export Demand On Sectoral and Main Trading Partners Classification

In this section, for the six countries given in Section 1 and for some sectors, which are important for Turkish economy, foreign income, production and real exchange rate elasticities of export demand are calculated. On the estimation of a group of related variables, it is meaningful to consider several models jointly. The classical panel data analysis investigates only the intercept difference across individuals or time periods whereas using the SUR estimation method the differences of the slope coefficient between individuals can also be investigated. The model is given by:

$$Y_{i,t} = X_{i,t}\beta_i + \varepsilon_{i,t} \qquad \text{(i=1,2,...,N; t=1,2,...,T).} \tag{6.1}$$

$$Y_i = [Y_{i,1}, Y_{i,2}, ..., Y_{i,T}]', X_i = [X_{i,1}, X_{i,2}, ..., X_{i,T}]', \text{and } \varepsilon_i = [\varepsilon_{i1}, \varepsilon_{i2}, ..., \varepsilon_{iT}]' \tag{6.2}$$

The stacked N equations (each has T observations) system is shown like that:

$$Y = X\beta + \varepsilon\,. \tag{6.3}$$

$$E[\varepsilon] = 0\,,\; E[\varepsilon_i \varepsilon_j'] = \sigma_{ij} I_T$$

A set of regression equations qualifies as a seemingly unrelated system if the contemporaneous correlation among the disturbances of each pair of equations is zero. The hypothesis that a set of regression equations constitutes a SUR system can be tested by Breusch-Pagan Lagrange Multiplier statistic. The null hypothesis is:

H_0: the contemporaneous correlation among the disturbances of each pair of equations is zero. The test statistic is:

$$\lambda_{LM} = T \sum_{i=1}^{n-1} \sum_{j=i+1}^{n} r_{ij}^{\;2} \tag{6.4}$$

where r^2_{ij} is the correlation between the residuals of *i*th and *j*th equations, T is the sample size and n is the number of equations in the system. Under the null hypothesis, the LM test statistic is distributed asymptotically chi-square with degrees of freedom n(n-1)/2.

The estimators that have been developed for simultaneous equations models are all based on the instrumental variable estimators. They differ in the choice of instruments and in whether the equations are estimated one at a time or jointly. The equations can be estimated one at a time using limited information methods or the equations can be estimated jointly using full information methods. Estimation of the system one equation at a time brings computational ease whereas joint estimation brings efficiency gains.[5]

In the literature, the use of limited information and full information methods are widely discussed. The general opinion is that, full information methods (3SLS and FIML) give better test results than limited information methods (2SLS and LIML) and have to be preferred. In this study three-stage least squares (3SLS) is used as estimation method.

6.1 Price and Production Elasticities of Sectoral Export Demand

In this section, real exchange rate and domestic production elasticities of the sectoral exports are calculated. The sectors under investigation are constructed according to the international standard industry classification (ISIC 3). The list of the sectors are given below:

-Food Products and Beverages
-Textiles
-Coke, Petroleum Products and Nuclear Fuel
-Chemicals and Chemical Products
-Manufacture of Basic Metals
-Manufacture of Machinery and Equipment
-Electrical Machinery and Apparatus
-Motor Vehicles and Trailers
-Mining and Quarrying

Initially, the LM test statistic (6.4) is calculated to see whether the nine equations constitute a seemingly unrelated regression system. The null hypothesis of zero contemporaneous correlation is rejected, which indicates that the nine equations constitute a SUR system. The model is given as:

$$\Delta \log y_{i,t} = \alpha_i + \lambda_{1i} S_1 + \lambda_{2i} S_2 + \lambda_{3i} S_3 + \beta_{1i} \Delta \log x_{1i,t} + \beta_{2i} \Delta \log x_{2i,t} + \varepsilon_{i,t} \quad (6.1.1)$$

$$i=1,\ldots,9 \; ; \; t=1,\ldots,25$$

The model contains individual specific intercept and seasonal dummies. The stationarity of series is investigated through ADF unit root test. Although the results

[5] Greene, W. H., (1997-3rd edition): Econometric Analysis, Prentice-Hall International, Inc.

are not given here, all the series are found to be I(1) and first difference transformation is done in order to obtain stationarity.

The variables used are:

$y_{i,t}$: sectoral export deflated by export price index

$x_{1i,t}$: sectoral industrial production index

$x_{2i,t}$: real exchange rate

$$x_{2i,t} = \sum_{j=1}^{6} w_j \frac{P_i}{e_j * P_j} \text{, j=1,…,6 (countries), i=1,…,9 (sectors)} \qquad (6.1.2)$$

where $w_j =$ (exports to country j) / (total exports of Turkey)

-e denotes the nominal exchange rate of the six countries

- P_i is Wholesale Price Index for sector i (the subscript d denotes domestic)

- P_j is the producer prices of country j (the subscript f denotes foreign).

The model estimations are based on quarterly data between the years 1994-2000 and the base year is changed to 1995. In Table 10, the calculated elasticities are given.

The real exchange rate elasticity of export demand is less than one for food, textiles, machinery and chemicals sectors whereas it is greater than one for remaining sectors. The sectoral production elasticities of textiles and machinery sectors are also smaller compared to other sectors. When the exported goods of food and textiles sectors are assumed to be classified as export of consumption goods, similar results with Vehbi (2002) can be reached. The author concluded that the foreign demand elasticity of export of consumption goods is close to one whereas its price elasticity is less than one. Bearing in mind that Turkish export is concentrated especially on labor-intensive sectors with low foreign demand like textiles, this portrait indicates a structural problem in the export.

The real exchange rate and domestic production elasticities of basic metals, motor vehicles, electrical machinery, coke, petroleum and mining sectors are greater than one. One of the interesting results obtained from empirical analysis is the high sectoral production elasticity of export demand for coke, petroleum and mining sectors. It is known that, the prices of most of the metals are determined in the international stock exchanges. In this respect, the sensitivity of coke, petroleum and mining sectors export demand to price changes is not surprising.

The sign of the real exchange rate elasticity is different among sectors. According to the real exchange rate definition given above a negative sign is expected. It means a depreciation of the Turkish lira or an increase in the foreign prices will cause an increase in the export volume. But the elasticities obtained from the regression equations for sectors related mostly with unprocessed and semi-processed primary goods such as basic metals, food, chemicals and mining are conflicting with the expected sign. An explanation of these positive elasticities can be the domestic demand expansions occurring as a result of the appreciation of Turkish lira. Another explanation can be an increase in the production costs of these sectors, which may cause an increase in the domestic prices and as a result of this a decrease in the domestic demand. This may direct exporters to foreign markets.

6.2 Income and Price Elasticities of Export Demand on Main Trading Partners Classification

In this part of the study, income and price elasticities of export demand are calculated on the basis of countries given in the Section 1.

At first, the LM test statistic given in (6.4) is calculated for testing whether the six equations constitute a seemingly unrelated regression system. The null hypothesis of zero contemporaneous correlation is rejected, which means that the six equations constitute a SUR system. The model estimated is given like that:

$$\Delta \log y_{i,t} = \alpha_i + \beta_{1i} \Delta \log X_{1i,t} + \beta_{2i} \Delta \log X_{2i,t} + \varepsilon_{i,t} \qquad (6.2.1)$$

i=1,…,6 ; t=1,…,47

The variables are:

$y_{i,t}$: real export to six countries

$x_{1i,t}$: the volume index of Gross Domestic Product (1995=100) series of country i

$x_{2i,t}$: is the real exchange rate of country i given by the formula

$$x_{2i,t} = \frac{P_d}{e * P_i} \qquad (6.2.2)$$

where,

- e_i denotes the nominal exchange rate of the six countries given above
- P_d is the domestic wholesale price index
- P_i is the producer price index of country i.

The model estimations are based on quarterly data between the years 1989-2000 and the base year is changed to 1987. In Table 11, the calculated elasticities are given.

As it can be seen from Table 11, the foreign income elasticities are found insignificant for Germany, France and Italy. For all remaining countries the foreign income elasticities are found to be significant and positive. The real exchange rate elasticities have negative sign for all countries as expected. Except Italy, the elasticites of real exchange rate for exports are less than one. The country base export demand analysis is in harmony with the aggregate export demand estimation.

7. Conclusions

In this paper, the export demand elasticities of foreign income, real exchange rate and sectoral production are estimated by using cross-section data. One of the objectives of this paper is the investigation and application of some panel data methods.

In empirical analysis, LL and IPS panel unit root tests and several cointegration tests including Pedroni panel cointegration test are performed. By finding evidence in favor of the cointegration relationship between variables, an error correction model is estimated for total exports by using the data of the six trade partners of Turkey. The export performance is also investigated on the sectoral and main trading partners bases. The conclusions emerging from the empirical results may be summarized as follows:

Aggregate export demand is found to be foreign income elastic both in the long-run and in the short-run. This can be interpreted as growth in trade partner countries may affect Turkey's export positively and significantly. But aggregate export demand is found to be real exchange rate inelastic both in the long-run and short-run. This gives support to the hypothesis that the exchange rate policies may not be successful in promoting export growth**.** The results obtained from SUR system of main trading partners also signal the significance of foreign demand. The real exchange rate elasticities obtained from SUR system of main trade partners are consistent with the results obtained from aggregate estimation of export demand. In a developing country like Turkey, export growth may be more dependent to factors like foreign demand, production capacity, productivity, diversification of exported goods and production of technology-intensive goods rather than price changes. In this respect, low price elasticity may not be surprising.

The results driven from sectoral base analysis indicate that exports react to both the sectoral production and real exchange rate. The real exchange rate elasticity of export demand is less than one for food, textiles, machinery and chemicals sectors, whereas it is greater than one for remaining sectors. The production elasticities of sectoral export demand are greater than one for all sectors except machinery sector. One interesting result obtained from the empirical analysis is the high sectoral production elasticity of export demand for coke, petroleum and mining sectors. The differences among the elasticities obtained from aggregated export demand and sector specific export demand estimation show the significance of investigating the dynamics of export on sectoral disaggregation.

To conclude it can be said that, effects of exchange rate policies on exports seems to be fairly limited. In order to obtain a sustainable and stabilized export growth, trade policies, which are based on diversification of exported products and production of technology-intensive goods, have to be developed.

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Table 10 The Real Exchange Rate and Domestic Production Elasticities of Export Based on Selected Sectors

	Basic Metals			Food			Textiles		
Dependent variable: Δlnex									
Independent variables	***Coefficient***	***t-statistics***	***p-value***	***Coefficient***	***t-statistics***	***p-value***	***Coefficient***	***t-statistics***	***p-value***
constant	0.175	0.134	0.894	-0.121	-1.310	0.194	-0.156	-2.139	0.035
s1	0.055	0.513	0.609	-0.403	-6.302	0.000	0.152	2.400	0.019
s2	0.592	4.940	0.000	0.310	1.490	0.140	0.341	3.727	0.000
s3	-0.380	-2.563	0.012	0.765	4.959	0.000	0.112	1.616	0.110
Δlnreer	1.275	2.949	0.004	-	-	-	-0.672	-2.306	0.024
Δlnreer(-1)	-	-	-	-	-	-	-	-	-
Δlnreer(-2)	-	-	-	0.656	1.800	0.075	-	-	-
Δlnreer(-3)	-	-	-	-	-	-	-	-	-
Δlnreer(-4)	-	-	-	-	-	-	-	-	-
Δlnprod	-	-	-	-	-	-	-	-	-
Δlnprod(-1)	2.838	5.098	0.000	1.281	3.200	0.002	1.229	3.370	0.001
Δlnprod(-2)	-	-	-	2.344	7.351	0.000	-	-	-
Δlnprod(-3)	-	-	-	-	-	-	-1.026	-3.948	0.000
Δlnprod(-4)	-	-	-	-	-	-	-	-	-
Δlnex(-1)	0.089	0.595	0.554	-0.298	-2.561	0.012	-0.652	-4.375	0.000
Δlnex(-2)	0.834	5.586	0.000	-0.482	-3.264	0.002	-0.766	-4.251	0.000
Δlnex(-3)	0.068	0.472	0.638	-0.846	-6.652	0.000	-0.310	-2.784	0.007
Δlnex(-4)	-	-	-	-	-	-	-	-	-
Diagnostics									
			p-value			***p-value***			***p-value***
R-squared	0.632			0.876			0.560		
Durbin-Watson statistic	2.476			2.112			2.063		
Jarque-Bera	2.054		0.358	1.425		0.490	0.308		0.857
Serial LM	6.792		0.147[1]	5.072		0.167[2]	4.624		0.328[1]

[1] Serial correlation LM test up to lag order 4.
[2] Serial correlation LM test up to lag order 3.
-Δ denotes first difference, log represents the logarithm transformation.

Table 10 continued

Dependent variable: Δlnex	Motor Vehicles			Elec. Machinery			Machinery		
Independent variables	***Coefficient***	***t-statistics***	***p-value***	***Coefficient***	***t-statistics***	***p-value***	***Coefficient***	***t-statistics***	***p-value***
constant	1.462	3.222	0.002	-0.002	-0.015	0.988	0.271	3.113	0.003
s1	-1.206	-3.069	0.003	-0.170	-2.053	0.043	-0.135	-1.584	0.117
s2	-0.988	-4.023	0.000	0.202	1.955	0.054	0.136	1.499	0.138
s3	-1.880	-4.187	0.000	0.292	1.886	0.063	-0.295	-3.448	0.001
Δlnreer	-	-	-	-	-	-	0.753	1.711	0.091
Δlnreer(-1)	7.109	2.630	0.010	-	-	-	-	-	-
Δlnreer(-2)	-	-	-	-2.808	-2.337	0.022	-	-	-
Δlnreer(-3)	-4.072	-2.320	0.023	2.984	3.208	0.002	-	-	-
Δlnreer(-4)	-	-	-	-	-	-	-	-	-
Δlnprod	-	-	-	-	-	-	-	-	-
Δlnprod(-1)	-	-	-	1.953	5.573	0.000	0.845	2.766	0.007
Δlnprod(-2)	-	-	-	1.473	4.788	0.000	-	-	-
Δlnprod(-3)	-	-	-	0.566	2.248	0.027	-	-	-
Δlnprod(-4)	-	-	-	-	-	-	-	-	-
Δlnprod(-5)	1.727	2.284	0.025	-	-	-	-	-	-
Δlnex(-1)	-0.252	-1.396	0.166	-0.615	-4.634	0.000	-0.854	-5.143	0.000
Δlnex(-2)	-	-	-	-0.658	-3.709	0.000	-0.429	-2.415	0.018
Δlnex(-3)	0.085	0.526	0.600	-0.602	-3.952	0.000	-	-	-
Δlnex(-4)	-	-	-	-	-	-	-	-	-
Diagnostics									
			p-value			***p-value***			***p-value***
R-squared	0.678			0.773			0.324		
Durbin-Watson statistic	2.722			1.982			2.027		
Jarque-Bera	0.315		0.854	1.807		0.405	0.570		0.752
Serial LM	7.039		0.134[1]	3.839		0.147[2]	7.568		0.109[1]

[1] Serial correlation LM test up to lag order 4.
[2] Serial correlation LM test up to lag order 4.
-Δ denotes first difference, log represents the logarithm transformation.

Table 10 continued

Dependent variable: Δlnex	Coke, petroleum			Chemicals			Mining		
Independent variables	***Coefficient***	***t-statistics***	***p-value***	***Coefficient***	***t-statistics***	***p-value***	***Coefficient***	***t-statistics***	***p-value***
constant	0.885	2.597	0.011	0.264	2.765	0.007	0.006	0.026	0.979
s1	-0.159	-0.513	0.609	0.119	1.437	0.154	2.003	5.192	0.000
s2	-1.172	-2.645	0.010	-0.100	-1.118	0.266	0.233	0.789	0.432
s3	-1.050	-2.902	0.005	-0.233	-1.773	0.080	-1.548	-5.101	0.000
Δlnreer	-	-	-	0.748	1.681	0.096	-	-	-
Δlnreer(-1)	-	-	-	0.824	1.878	0.064	3.524	3.901	0.000
Δlnreer(-2)	2.650	3.118	0.003	-	-	-	-	-	-
Δlnreer(-3)	-	-	-	-	-	-	-2.436	-2.668	0.009
Δlnreer(-4)	-	-	-	-	-	-	-	-	-
Δlnprod	9.894	5.414	0.000	-	-	-	6.046	5.687	0.000
Δlnprod(-1)	-	-	-	2.417	3.701	0.000	5.091	5.229	0.000
Δlnprod(-2)	-	-	-	-	-	-	-	-	-
Δlnprod(-3)	9.186	6.443	0.000	-	-	-	-	-	-
Δlnprod(-4)	-	-	-	-	-	-	-	-	-
Δlnex(-1)	-0.373	-3.645	0.000	-0.473	-3.106	0.003	0.133	1.025	0.308
Δlnex(-2)	-	-	-	-0.785	-4.166	0.000	0.583	3.623	0.001
Δlnex(-3)	-	-	-	-0.392	-2.357	0.021	-	-	-
Δlnex(-4)	-	-	-	-0.364	-2.107	0.038	-0.325	-2.634	0.010
Δlnex(-5)	-	-	-	-0.841	-4.710	0.000	-	-	-
Diagnostics									
			p-value			***p-value***			***p-value***
R-squared	0.786			0.503			0.614		
Durbin-Watson statistic	1.805			1.948			1.729		
Jarque-Bera	1.846		0.397	1.238		0.538	0.356		0.837
Serial LM	3.965		0.411[1]	7.142		0.129[1]	5.632		0.131[2]

[1] Serial correlation LM test up to lag order 4.
[2] Serial correlation LM test up to lag order 3.
-Δ denotes first difference, log represents the logarithm transformation.

Table 11 The Real Exchange Rate and Foreign Income Elasticities of Export Based on Selected Countries

	USA			Germany			France		
Dependent variable: Δlnex									
Independent variables	***Coefficient***	***t-statistics***	***p-value***	***Coefficient***	***t-statistics***	***p-value***	***Coefficient***	***t-statistics***	***p-value***
constant	-0.050	-1.408	0.161	0.056	4.165	0.000	0.064	3.507	0.001
Δlnincome	-	-	-	-	-	-	-	-	-
Δlnincome(-1)	14.602	3.884	0.000	-	-	-	-	-	-
Δlnincome(-2)	-	-	-	-	-	-	-	-	-
Δlnincome(-3)	-	-	-	-	-	-	-	-	-
Δlnincome(-4)	-	-	-	-	-	-	-	-	-
Δlnreer	-	-	-	-	-	-	-	-	-
Δlnreer(-1)	-0.732	-1.891	0.060	-	-	-	-0.658	-2.982	0.003
Δlnreer(-2)	-1.364	-3.754	0.000	-	-	-	-	-	-
Δlnreer(-3)	-	-	-	-0.340	-1.803	0.073	0.555	2.339	0.020
Δlnreer(-4)	-	-	-	-	-	-	-	-	-
Δlnex(-1)	-0.476	-4.372	0.000	-0.478	-3.057	0.003	-0.544	-4.810	0.000
Δlnex(-2)	-0.673	-8.406	0.000	-0.904	-11.138	0.000	-0.600	-4.653	0.000
Δlnex(-3)	-0.436	-3.996	0.000	-0.788	-3.414	0.001	-0.425	-3.597	0.000
Δlnex(-4)	-	-	-	-0.413	-1.804	0.073	-0.193	-1.898	0.059
Diagnostics									
			p-value			***p-value***			***p-value***
R-squared	0.735			0.784			0.509		
Adjusted R-squared	0.690			0.744			0.424		
Durbin-Watson statistic	2.193			2.393			2.063		
Jarque-Bera	0.498		0.779	0.736		0.692	1.435		0.488
Serial LM	4.217		0.377[1]	4.758		0.093[2]	2.287		0.683[1]

[1] Serial correlation LM test up to lag order 4.
[2] Serial correlation LM test up to lag order 2.
-Δ denotes first difference, log represents the logarithm transformation.

Table 11 continued

	Italy			Netherlands			UK		
Dependent variable: Δlnex									
Independent variables	***Coefficient***	***t-statistics***	***p-value***	***Coefficient***	***t-statistics***	***p-value***	***Coefficient***	***t-statistics***	***p-value***
constant	0.020	0.933	0.352	0.015	0.382	0.703	0.053	3.207	0.002
Δlnincome	-	-	-	8.911	4.697	0.000	-	-	-
Δlnincome(-1)	-	-	-	-	-	-	-	-	-
Δlnincome(-2)	-	-	-	-	-	-	5.724	2.636	0.009
Δlnincome(-3)	-	-	-	-	-	-	-	-	-
Δlnincome(-4)	-	-	-	-	-	-	-	-	-
Δlnreer	-1.133	-4.469	0.000	-	-	-	-	-	-
Δlnreer(-1)	-	-	-	-	-	-	-	-	-
Δlnreer(-2)	-	-	-	-	-	-	-0.372	-1.777	0.077
Δlnreer(-3)	-	-	-	-0.706	-2.001	0.047	0.503	2.379	0.018
Δlnreer(-4)	-	-	-	-	-	-	-	-	-
Δlnex(-1)	-0.182	-1.694	0.092	-0.890	-8.200	0.000	-0.657	-6.476	0.000
Δlnex(-2)	-0.205	-1.922	0.056	-0.982	-8.508	0.000	-0.642	-6.191	0.000
Δlnex(-3)	-0.289	-2.687	0.008	-0.896	-5.816	0.000	-0.510	-4.663	0.000
Δlnex(-4)	0.334	3.512	0.001	-	-	-	-	-	-
Diagnostics									
			p-value			***p-value***			***p-value***
R-squared	0.475			0.751			0.701		
Adjusted R-squared	0.402			0.716			0.650		
Durbin-Watson statistic	2.357			1.671			1.909		
Jarque-Bera	1.433		0.488	0.587		0.746	2.284		0.319
Serial LM	6.451		0.092[1]	5.411		0.144[1]	1.785		0.618[1]

[1] Serial correlation LM test up to lag order 3.

-Δ denotes first difference, log represents the logarithm transformation.

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