METU Studies in Development, 37 (April), 2010, 53-72

Testing for structural change in the bilateral trade elasticities of Turkey

İdil Uz
Department of Economics, Yeditepe University, Kayisdagi, 34755, Istanbul, Turkey
email: idiluz@yeditepe.edu.tr

Abstract

This paper estimates the bilateral trade equations of Turkey and its major trading countries by using the Gregory and Hansen procedure. We allowed for a structural break at an unknown date within the Gregory and Hansen (GH) framework and obtained a cointegrating relationship especially between variables in the export and import demand equations at the bilateral level. The empirical analysis of the exchange rate and income elasticity of trade demand, over the period 1982-2007, is presented and discussed.

Keywords: Structural change; income elasticities; exchange rate elasticities; cointegration.

JEL: F11, F14.

1. Introduction

Trade elasticities have major macroeconomic policy implications for any country. The major determinants of international trade are the gross domestic product (GDP) of a country, the foreign GDP, the price of export and import, the foreign and domestic prices, and exchange rate. Since shifts in the trade balances are regarded widely as a function of changes in exchange rates, the relationship between trade and these determinants are examined by a robust model of the export and import demand functions (Haynes et al., 1996).

High implied values of the trade elasticity are related to Turkey's connection to international prices and real income values. On the other hand, low values are related to disconnection from international prices and real income variables. Both cases provide valuable information for firms and traders to consider such relationships when contemplating changes in price, income, demand, and supply. The elasticity approach may give potential explanations of the behaviour of current account balance and also the consumers' preferences on domestic and imported goods. The elasticity approach helps to permit a distinct investigation of risk and expectation effects on exports and imports. Furthermore, the value of the elasticity of substitution determines how much a monetary authority should care about exchange rate movements. The position of the current account balance and its effect on exchange rates is an important indicator for the central bank in determining interest rates. In other words, if domestic and foreign goods are close substitutes, a central bank should not care very much about the exchange rate, because shocks to the domestic economy will not affect international trade much. Then protectionist defences against the international diffusion of recessions seem out of order. In general, the magnitude of the elasticity of substitution will substantially affect the policy implications of most models in international economics, beyond monetary or trade policy.

This study differs from earlier work by including real trade values and real exchange rates. This reduces the possibility of the bias caused by omitting a potentially relevant variable from the estimation. This study also differs from earlier work by examining trade elasticities on a bilateral basis (including Turkey's major trading partners, such as Canada, China, France, Germany, Italy, Japan, South Korea, the Netherlands, Russia, Spain, Switzerland, the UK, and the US), rather than aggregated trade and price data. This eliminates many of the problems associated with defining and using aggregate variables. Additionally, following Togan and Berument (2007), the prices of exports and imports are used in calculating the real import and export functions, alongside the real exchange rates. Furthermore, the implication of trade function allows for testing whether the Marshal Lerner condition holds for bilateral trade in Turkey in the long run.

Testing for cointegration between variables with unit roots is an integral part of empirical time series analyses. A number of tests are available in the literature. Most of these tests are residual-based and they are widely used due to their simplicity. However, these tests were introduced based on the assumption that the cointegrating vector remained the same during the period of study. There are many reasons to expect that the long-run relationship between the underlying variables might change. Structural changes can take place because of economic crises, technological shocks, changes in the economic actors' preferences and behaviour accordingly, policy and regime changes, and organizational or institutional evolution. This is more likely to be the case if the time span is long. This study covers the period between 1982Q2 and 2007Q4. 1980 constitutes a turning point in the economic history of Turkey. The import substitution strategy and the fixed exchange rate policy of the previous period were replaced with the export promotion strategy and the floating exchange rate policy. The establishment of a free market economy and integration with the world economy became the major objectives of the national policies, which were also supported by international organisations such as the IMF and the World Bank. The time span included in this study covers the transition period of the trade liberalisation started with the 1980s liberalisation policies and the Customs Union (CU) Agreement in 1996 between Turkey and the European Union (EU). These were considered as the major important factors that affected the Turkish economy. Nevertheless, there was a gradual change from protectionist policies towards liberalisation. Trade relations with the EU were based on a progressive establishment of the CU that took its origins from the Ankara Agreement, signed on 12 September 1963, and started with the Additional Protocol, which came into force in 1973. Therefore, in analysing the long-run behaviour of export and import demand functions for Turkey, one may expect a regime shift in the economy activity that occurred at some unknown date. The purpose of this study and its contribution to related literature is to test for a structural change in the long-run relationship of bilateral trade elasticities in Turkey.

The empirical analysis of the exchange rate and the income elasticity of trade demand covers the period 1982-2007. The empirical findings of this paper support the findings of Neyaptı et al. (2007). In the long run, exchange rate is elastic at the bilateral level for Germany in export demand function and for Canada, China, Spain, the Netherlands and Spain in import demand function. This shows Turkey's connection to prices in those countries. Furthermore, foreign demand is sensitive to exports and domestic demand is sensitive to imports. However, the short-run behaviour of real exchange rate to long-run disequibrium is statistically significant in all countries other than Russia, but small in values. Nevertheless, we found no evidence that trade is income elastic in the short run.

Finally, the study is structured as follows: Section 2 gives the literature review, section 3 gives the theoretical framework, section 4 explains the methodology, section 5 gives the unit root and cointegration test results. In section 6, data and the empirical findings are explained. Finally, section 7 makes the conclusion.

2. Literature review

A number of studies have followed the traditional approach and have estimated import and export demand elasticities to determine whether the Marshal-Lerner (ML thereafter) condition holds. See, for example, Kreinin (1967), Houthakker and Magee (1969), Khan (1974, 1975), Goldstein and Khan (1976, 1978), Wilson and Takacs (1979), Haynes and Stone (1983), Warner and Kreinin (1983), and Bahmani-Oskooee (1986). The ML condition states that as long as the sum of the price elasticity of export and import demand functions exceed unity, devaluation will improve the trade balance. Additionally, there have been studies on estimating trade elasticities for developing countries. Bahmani-Oskooee and Niroomand (1998) tested long run price elasticities and Marshal-Lerner condition for thirty developed and developing countries. Lal and Lowinger (2002) confirmed the existence of both short-run and long-run relationships between nominal exchange rate and trade balances for South Asia countries.

One of the criticisms of these studies has been the use of aggregate trade data. The problem, so-called "aggregation bias," is that a significant price elasticity with one trading partner could be more than offset by an insignificant elasticity (see Bahmani-Oskooee and Goswami, 2004). Therefore, this opens a new research area for the study of trade elasticities on a bilateral basis.

Some studies include bilateral trade between selected developed countries and different regions such as that of Marquez (1990). There have been studies on the bilateral trade between the US and one or more of its trading partners, for example, Cushman (1990), Haynes et al. (1996), Bahmani-Oskooee and Brooks (1999), and Nadenichek (2000). However, there also have been studies on the bilateral trade of one country other than the US; for example, studies of bilateral trade in Canada by Bahmani-Oskooee et al. (2005), or bilateral trade in Sweden by Hatemi-J and Irandoust (2005), Irandaust et al. (2006) and of the bilateral trade of manufacturing goods in Japan by Harriigan and Vanjani (2003). Nevertheless, even fewer studies have been conducted for the analysis of bilateral trade in developing countries beside the studies by Wang and Ji (2006) for bilateral trade in China and Liu et al. (2007) for bilateral trade in Hong Kong. This study aims to fill this gap and study bilateral trade elasticity between Turkey and its major trading partners.

There have been studies on the trade elasticity of merchandise imports and exports for the Turkish economy. The relation between Turkey's export and exchange rate is controversial. For example, Neyaptı et al. (2007) examined the effects of the Customs Union Agreement between Turkey and the EU on the behaviour of Turkey's export and import demands. In other studies, Celasun and Rodrik (1989) found little support for establishing a relationship between the export and exchange rate policy, where Arslan and Wijnbergen (1993) found a positive relationship between exports and the domestic depreciation of the currency. Nevertheless, there are studies showing that the Marshal–Lerner condition holds as the absolute values of estimated price elasticities for the import and export of goods add up to more than unity (Şimşek and Kadılar, 2004; Togan and Berument, 2007).

For the income elasticity of merchandise exports, Aydin et al. (2004) found a positive relation between elastic income in the long run and inelastic income in the short run. Özkale and Karaman (2006) showed that Turkey's import demand is income elastic and price inelastic. However, Özkale and Karaman (2006) found inelastic income elasticity for the period after the establishment of the Customs Union between Turkey and the EU (for the period between 1996 and 2004), with a negative sign for the income elasticity coefficient. Şimşek and Kadılar (2004) found that there is a long-run relationship between the export or import of goods and price and income. Aydın et al (2004) and another recent study by Togan and Berument (2007) found elastic foreign demand for the export of goods with a positive sign.

Stability of export and import demand coefficients was the centre of attention in studies using aggregate trade data. For example, Doğanlar (1998) argued that after 1980, Turkish exports and terms of trade followed a different growth path while Turkish imports did not change its growth path. In another study Utkulu and Seymen (2004) found long-run cointegration relationship on aggregate level among the variables in export and import demand equations, dealing with the hidden structural breaks. They concluded that dummies for structural breaks lower both the long-run price and income elasticities for Turkish exports with the EU, where dummies lower only price elasticity for Turkish imports from the EU. Thomakos and Ulubaşoğlu (2002) estimated disaggregated import demand elasticities for Turkey for various product groups over the period 1970-1995. They have tested for different elasticities and found that the effects of the trade reforms of the 1980's were significant for a number of industries that form the backbone of the Turkish economy.

3. Theoretical framework

The effect of income and the real exchange rate on international trade is well recognized in the literature. To examine to what extent movements in the balance of the trade of services are explained by change in relative prices, income and exchange rate we employ an imperfect substitute model (Goldstein and Khan, 1985) for the export and import demand function, where we assume that foreign and domestic products are imperfect substitutes.

$$X_{it} = \mathrm{f}(P_{xt}, P_{it}^*, Y_{it}^*) \tag{1}$$

where t denotes the time period of estimation, X_{it} is the value of the export of goods to ith country, P_x is the export price in New Turkish Lira (domestic GDP deflator is taken as a proxy), P_i^* denotes the foreign price deflator (GDP deflator is used as a proxy) in New Turkish Lira, and Y_i^* is foreign real GDP. If we divide the right-hand side of the equation (1) by foreign prices P_{it}^*, due to the linearity of demand functions the export demand is not going to change (Goldstein and Khan, 1985). Therefore the logarithmic form of the export demand function is as follows:

$$\mathrm{lnRX}_{\mathrm{it}} = \alpha_0 + \alpha_1 \ln q_{it} + \alpha_2 \ln Y_{it}^* + D1 + D2 + D3 + \varepsilon_{\mathrm{t}} \tag{2}$$

where $\mathrm{lnRX}_{\mathrm{it}}$ is the natural log of real export value calculated by export from the ith country deflated by price of exports. $\ln q_{it}$ is the natural log of real exchange rate calculated by $\ln(P_{it}^* e/P_t)$, where P_{it}^* is foreign price, e is nominal exchange rate and P_t is domestic price level and $\ln Y_{it}^*$ is the natural log of the foreign real income. D1 is the dummy variable associated with switching to euro for euro zone countries, D2 is the dummy variable associated with the Asian and Russian crises and D3 is the dummy variable associated with the financial crises in Turkey[1]. Finally ε_t is the error term.

We expect a coefficient of relative price (real exchange rate) α_1 in equation 2 to be positively related to export because an increase in the real exchange rate, a depreciation of the Turkish Lira (TL) promotes Turkish competitiveness and thus its exports. It is difficult to define the sign of income elasticity α_2 as it can have a different sign. From one side increase in the foreign income can raise demand for Turkish exports. However, if foreign goods are highly competitive with Turkish exports foreign income in this case can have a negative effect on the export volume from Turkey.

The standard form of the import demand function can be expressed by the following equation:

$$M_{it} = f(P_{mt}, P_t, Y_t) \tag{3}$$

where M_{it} is the value of import from the ith country, P_{mit} denotes the import price of the traded goods in New Turkish Lira (the foreign GDP deflator is taken as a proxy), P_t denotes a domestic price deflator (the domestic GDP deflator is used as a proxy) and Y_t is domestic real GDP. Following the analysis made in export demand function extraction we can divide the right-hand side of equation (3) by domestic prices P_t. As a result, the estimated import demand function is as follows:

$$\ln \mathrm{RM}_{it} = \beta_0 + \beta_1 \ln q_{it} + \beta_2 \ln Y_t + D1 + D2 + D3 + \upsilon_t \tag{4}$$

where $\ln \mathrm{RM}_{it}$ is the natural log of real import quantity deflated by the price of imports and $\ln q_{it}$ is the natural log of the real exchange rate. β_1 is expected to have a negative sign because real depreciation is expected to reduce Turkish imports. $\ln Y_t$ is the natural log of the domestic real income. If the income elasticity of β_2 is positive, this implies that Turkish income leads to an increase in Turkish imports. D1, D2 and D3 are dummy variables explained in the export demand function. Finally υ_t is the error term.

We assume that the relative import prices coefficients β_1 will be negatively related to the import quantity as according to the demand theory an increase in the import price will reduce the import demand while an increase in domestic prices will raise demand for import. Income coefficient, β_2 , is expected to have a positive sign in most of the cases but as well as it may have a negative effect on import demand. If there are no any alternatives for imported goods in the domestic production, income will have a positive effect on the import volume.

[1] See appendix 1 for further details of the dummy variables.

4. Methodology

The first consideration is the stationarity or otherwise of the time series employed in this analysis. Standard unit root tests including the Dickey and Fuller (1979) Augmented Dickey-Fuller (ADF) test, the Phillips-Perron (1988) test (PP), and Kwiatkowski, Phillips, Schmidt, and Shin's (1992) KPSS test results for the individual time series are conducted for this study[2]. However, since the aim of this analysis is to examine the long-term behaviour of bilateral trade under regime shift, we follow the procedure suggested by Perron (1989) and Zivot and Andrews (1992). It is well known that structural breaks in the deterministic components of the stochastic process tend to bias both ADF and PP tests towards the unit root null hypothesis; for this reason Perron (1989) proposes both innovative and additive type outlier test of a I(1) null hypothesis (non-stationary) against a I(0) alternative (stationary) with a single break, occurred at an unknown point in time denoted as T_b. Unit root null hypothesis are developed under three models that are change in level, change in slope of linear trend and change in level and slope of linear trend.

$$\text{Model (A)} \qquad y_t = \mu_1 + \beta t + (\mu_2 - \mu_1)DU_t + e_1$$

$$\text{Model (B)} \qquad y_t = \mu_1 + \beta_1 t + (\beta_2 - \beta_1)DT_t^* + e_1$$

$$\text{Model (C)} \qquad y_t = \mu_1 + \beta_1 t + (\mu_2 - \mu_1)DU_t + (\beta_2 - \beta_1)DT_t^* + e_1$$

where DU_t =1, DT^* = T-T_b if $t > T_b$ and 0 otherwise. Model A allows for a one-time change in the level of the series and the difference $(\mu_2 - \mu_1)$ represents the magnitude of the change in the intercept of the trend function occurring at time T_b. In Model B, the difference $(\beta_2 - \beta_1)$ represents the magnitude of the change in the slope of the trend function occurring at time T_b. Model C combines the change in the level and the slope of the trend function of the series. Finally, T is the sample size and e_1 is the error term. When Perron's (1989) models are adapted to ADF type unit root test, it takes the following form:

$$\Delta y_t = \hat{\mu}^A + \hat{\theta}^A DU_t + \hat{\beta}^A t + \alpha^A \mathrm{y}_{\mathrm{t\text{-}1}} + \sum_{j=1}^{k} \hat{c}_j^A \Delta y_{t-1} + \hat{e}_{\mathrm{t}} \tag{5}$$

$$\Delta y_t = \hat{\mu}^B + \hat{\beta}^B t + \hat{\gamma}^B DT_t^* + \alpha^B \mathrm{y}_{\mathrm{t\text{-}1}} + \sum_{j=1}^{k} \hat{c}_j^B \Delta y_{t-1} + \hat{e}_{\mathrm{t}} \tag{6}$$

[2] Since the aim of this study is to test for a structural break in the cointegration analysis, we did not give ADF, PP and KPSSS test results in the text, but they are available upon request.

$$\Delta y_t = \hat{\mu}^C + \hat{\theta}^C DU_t + \hat{\beta}^C t + \hat{\gamma}^C DT_t^* + \alpha^C \mathrm{y}_{\mathrm{t\text{-}1}} + \sum_{j=1}^{k} \hat{c}_j^C \Delta y_{t-1} + \hat{e}_{\mathrm{t}} \tag{7}$$

equations 5, 6 and 7 above correspond to Model A, B and C, respectively. *k* is the number of autoregressive lags used in the analysis. The appropriate lag length was set as in Perron (1997) following a general to specific recursive procedure based on the t-statistics on the coefficient associated with the last lag in the estimated auto-regression. *k* max is set by Schwert (1989)[3]. To test the presence of a unit root, t-statistic of the coefficient, α, is expected to be higher than the critical values to reject the null of unit root hypothesis. Critical values are reported by Zivot and Andrews (1992, pp. 256-57).

The power of the Engle-Granger (1987) (EG) test of the null of no cointegration is substantially reduced when there is a break in the cointegrating relationship. To overcome this problem, Gregory and Hansen (1996a) (GH) extend the EG test to allow for breaks in either the intercept or the intercept and trend of the cointegrating relationship at an unknown time. Gregory and Hansen (1996a, 1996b) developed 4 models that are used in this analysis. These are the level shift model (C), the level shift model with trend (C/T), the regime shift model (C/S) and the regime and trend shift model (C/S/T).

(C) $y_{1t} = \mu_1 + \mu_2 DU_t + \alpha \mathrm{y}_{2\mathrm{t}} + e_{\mathrm{t}}$

(C/T) $y_{1t} = \mu_1 + \mu_2 DU_t + \beta t + \alpha \mathrm{y}_{2\mathrm{t}} + e_{\mathrm{t}}$

(C/S) $y_{1t} = \mu_1 + \mu_2 DU_t + \alpha_1 \mathrm{y}_{2\mathrm{t}} + \alpha_2 DUy_{2t} + e_{\mathrm{t}}$

(C/S/T) $y_{1t} = \mu_1 + \mu_2 DU_t + \beta_1 t + \beta_2 DT_t^* \alpha_1 \mathrm{y}_{2\mathrm{t}} + \alpha_2 DUy_{2t} + e_{\mathrm{t}}$

In the above parameterisations μ_1 represents the intercept before the shift and μ_2 represents the change in the intercept at the time of the shift. Similarly, β_1 represents the time trend before the shift and β_2 represents the time trend at the time of the shift. α_1 denotes the cointegrating slope coefficients before the regime shift, and α_2 denotes the change in the slope coefficients.

GH modified three residual based unit root tests that take into account one unknown regime shift. They suggest a procedure to choose the timing of a shift in the cointegrating vector based on the data. They furthermore provide new critical values for the ADF test for cointegration (as suggested

[3] In this study, following Schwert (1989), k max is set as 12 and $k\ \max = \left[12.\left(\frac{T}{100} \right)^{1/4} \right]$

by EG, 1987). The current paper builds on the GH tests for cointegration in the presence of one shift at an unknown date, t_b. The ADF test results are reported in section 5.2 for each breakpoint in the interval, 0.15T to 0.85T (where T is the number of observations).

Several of the variables used for testing trade elasticities were non-stationary, making the use of standard econometric procedure such as ordinary least squares inappropriate. Therefore, cointegration analysis will be used for testing the long-run relationship in export and import demand functions. It is common in the literature to use various cointegration techniques to offset the disadvantages of each technique (see, for example, Chinn, 2005; Narayan and Narayan, 2005; Makin and Narayan, 2008; Uz and Ketenci, 2008). For estimating a long-run (cointegrating) relationship between the variables, it is common in literature to employ techniques such as the dynamic OLS (DOLS) estimator by Kao and Chiang (2000), Johansen's (1988) multivariate maximum-likelihood procedure (JOH-ML) with the vector error correction (VEC) framework that studies the short-run behaviour of trade demand functions. Kao and Chiang (2000) show that both the OLS estimator exhibit small-sample bias and that the DOLS estimator appears to outperform the OLS estimator. In order to obtain an unbiased estimator of the long-run parameters, the DOLS estimator uses parametric adjustment to the errors by augmenting the static regression with the leads, lags, and contemporaneous values of the regressors in first differences.

The deviation from long-run equilibrium corrected in real export and real import is tested for real exchange rate and income. A vector error correction model applied the Johansen's approach.

$$\Delta\omega_t = \sum_{j=1}^{k-1} \Gamma_j \Delta\omega_{t-j} + \Pi\omega_{t-k} + u_t \tag{8}$$

where ω_t is (nx1) vector, Γ_j and Π are (nxn) matrices of parameters representing short-term and long-run impacts, respectively. $\Pi = \alpha\beta'$, where α reflects the speed of adjustment toward equilibrium, while β is a matrix of long-run coefficients.

The short-run relationship in the export and import demand function is analysed by the vector error correction model. For the export demand function, the equation becomes as follows;

$$\Delta\ln\mathrm{RX}_t = \mu_1 + \sum_{i=1}^{p} \delta_{1i}\Delta\ln\mathrm{RX}_{t-i} + \sum_{i=1}^{p} \gamma_{1i}\Delta\ln q_{t-i} + \sum_{i=1}^{p} \kappa_{1i}\Delta\ln Y^*_{t-i} + \lambda_1\,\Omega + \omega_{1t}$$

$$\Delta\ln\ln q_t = \mu_2 + \sum_{i=1}^{p} \delta_{2i}\Delta\ln\mathrm{RX}_{t-i} + \sum_{i=1}^{p} \gamma_{2i}\Delta\ln q_{t-i} + \sum_{i=1}^{p} \kappa_{2i}\Delta\ln Y^*_{t-i} + \lambda_2\,\Omega + \omega_{2t}$$

$$\Delta\ln Y^*_t = \mu_3 + \sum_{i=1}^{p} \delta_{3i}\Delta\ln\mathrm{RX}_{t-i} + \sum_{i=1}^{p} \gamma_{3i}\Delta\ln q_{t-i} + \sum_{i=1}^{p} \kappa_{3i}\Delta\ln Y^*_{t-i} + \lambda_3\,\Omega + \omega_{3t} \tag{9}$$

where, Ω is the equilibrium relations that $\Omega = (\ln\mathrm{RX}_{t\text{-}1} - \ln q_{t-1} - \ln Y^{*}_{t-1})$, λ_1, λ_2 and λ_3 reflect the speed of adjustment toward the equilibrium.

For the import demand function, the equation becomes;

$$\Delta\ln\mathrm{RM}_t = \chi_1 + \sum_{i=1}^{p}\phi_{1i}\Delta\ln\mathrm{RM}_{t\text{-}i} + \sum_{i=1}^{p}\xi_{1i}\Delta\ln q_{t-i} + \sum_{i=1}^{p}\theta_{1i}\Delta\ln Y_{t-i} + \eta_1\,\Lambda + \omega_{1t}$$

$$\Delta\ln q_t = \chi_2 + \sum_{i=1}^{p}\phi_{2i}\Delta\ln\mathrm{RM}_{t\text{-}i} + \sum_{i=1}^{p}\xi_{2i}\Delta\ln q_{t-i} + \sum_{i=1}^{p}\theta_{2i}\Delta\ln Y_{t-i} + \eta_2\,\Lambda + \omega_{1t}$$

$$\Delta\ln Y_t = \chi_3 + \sum_{i=1}^{p}\phi_{3i}\Delta\ln\mathrm{RM}_{t\text{-}i} + \sum_{i=1}^{p}\xi_{3i}\Delta\ln q_{t-i} + \sum_{i=1}^{p}\theta_{3i}\Delta\ln Y_{t-i} + \eta_3\,\Lambda + \omega_{1t} \qquad (10)$$

where, Λ is the equilibrium relations that $\Lambda = (\ln\mathrm{RM}_{t\text{-}1} - \ln q_{t-1} - \ln Y_{t-1})$, η_1, η_2 and η_3 reflect the speed of adjustment toward the equilibrium. Additionally, a number of dummies were added to include the regime and trend shifts into the VEC model.

5. Unit root and cointegration tests

The analysis starts with investigating the integration properties of the variables necessary for estimating export and import demand models. The variables investigated for the unit root are real export, real import, real exchange rate and real GDP for Turkey's major trading partners: Canada, China, France, Germany, Italy, Japan, Korea, the Netherlands, Russia, Spain, Switzerland, the UK, and the US. The most controversial issue in selecting variables for cointegration is to identify the values with unit root. The cointegration test is employed only for variables with unit root.

Table 1 gives the unit root test results. T_α^* is the t-statistic for coefficient α in equations 5, 6 and 7. The test results show that all variables other than the real export for the Netherlands were found non-stationary at their level and stationary at their first differences. The empirical results for their first differences are not presented here out of space considerations. Therefore, real export demand for the Netherlands is excluded in the cointegration analysis.

The next step is to identify whether there is cointegration relationship between the real trade values and their determinants such as real exchange rate and income. The test results for the GH approach are presented in Table 2. The test results found some evidence of a long-run cointegrating relationship between variables in the export demand equation for European countries such as Germany, Spain, Switzerland and the UK and for variables in import demand equation for Canada, China, France, Germany, Japan, Netherlands, Russia, Spain, Switzerland and the US, after adopting a testing procedure that allows for an endogenous break identification.

Table1

Unit Root Test Results

	REAL EXPORTS						REAL IMPORTS					
	MODEL A		MODEL B		MODEL C		MODEL A		MODEL B		MODEL C	
	t_α^*	*k*	t_α^*	*k*	t_α^*	*k*	t_α^*	*k*	t_α^*	*k*	t_α^*	*k*
Canada	-2.35	7	-2.74	7	-2.66	7	-2.03	5	-2.62	5	-2.23	5
China	-2.62	2	-2.64	2	-2.56	2	-3.20	10	-4.43**	10	-1.73	10
UK	-2.64	7	-2.12	7	-1.19	7	-3.76	4	-3.29	4	-3.72	4
France	-3.17	8	-3.06	8	-4.02	8	-2.39	8	-2.39	8	-2.39	8
Germany	-3.10	10	-3.22	10	-2.83	10	-2.86	8	-2.61	8	-2.85	8
Italy	-2.26	10	-2.68	10	-2.92	10	-2.73	8	-2.41	8	-2.71	8
Japan	-2.65	5	-2.37	5	-4.58	5	-3.81	4	-3.64	4	-1.02	4
Korea	-2.34	6	-1.52	6	-2.64	6	-1.83	6	-2.34	6	-2.13	6
Netherlands	-4.80**	9	-5.03***	9	-5.74**	9	-2.74	3	-1.55	3	-2.66	3
Russia	-2.86	9	-2.88	9	-2.85	9	-4.07	4	-4.24*	4	-4.92*	4
Spain	-2.47	10	-2.51	10	-1.82	10	-3.48	1	-3.22	1	-3.86	1
Switzerland	-3.74	8	-4.09	10	-3.87	10	-3.77	8	-4.46**	8	-4.46	8
USA	-2.64	9	-2.82	9	-2.40	9	-0.48	7	-1.15	7	-0.83	7
	REAL EXCHANGE RATE						REAL INCOME					
Canada	-3.50	2	-3.59	2	-3.51	2	-2.87	3	-2.93	3	-2.93	3
China	-2.73	4	-2.78	4	-3.39	4	-1.17	8	-2.68	8	-2.54	8
UK	-5.59	6	-4.09	6	-5.59**	6	-2.65	3	-2.75	3	2.98	3
France	-3.58	10	-3.81	10	-3.19	10	-3.95	9	-3.64	9	-4.10	9
Germany	-2.74	7	-2.27	7	-1.66	7	-2.76	4	-1.64	4	-2.23	4
Italy	-3.38	5	-3.71	5	-4.48	0	-2.89	1	-1.16	1	-3.22	1
Japan	-1.87	0	-1.36	2	-0.23	0	-1.71	8	-0.47	8	-2.46	8
Korea	-3.09	2	-2.78	2	-3.13	2	-2.22	8	-0.79	8	-2.59	8
Netherlands	-3.22	1	-3.22	1	-1.61	1	-2.63	1	-2.83	1	-2.80	1
Russia	-2.47	10	-2.49	10	-2.40	10	-2.00	5	-1.78	5	-2.12	5
Spain	-2.29	0	-1.98	0	-2.82	0	-3.38	7	-3.68	7	-3.67	7
Switzerland	-3.17	3	-2.37	0	-3.33	3	-2.38	9	-2.77	9	-3.16	9
USA	-1.90	6	-2.65	6	-2.37	6	-3.35	9	-3.71	9	-3.07	9
Turkey	-	-	-	-	-	-	-2.38	8	-2.35	8	-3.19	8

Null of non-stationary is tested.

k is number of lags selected according to Perron(1997)

*, **, *** indicate significance at 10%, 5% and 1% levels, respectively.

Critical values are taken from Zivot and Andrews,1992, pp. 256-57.

Table 2
Cointegration Test Results

	EXPORTS				IMPORTS			
	C	C/T	C/S	C/S/T	C	C/T	C/S	C/S/T
Canada	-0.60	-0.47	-0.57	-0.76	-0.84***	-0.85***	0.43***	-0.87***
China	-0.23	-0.24	-0.41	-0.27	-0.40**	-0.40*	-0.59***	-0.65
France	-0.68	-0.72	-0.59	-0.65	-0.73**	-0.73	-0.75	-0.79
Germany	-0.49	-0.82**	-0.56	-0.95***	-0.44	-0.44	-0.43	-0.95***
Italy	-0.35	-0.27	-0.34	-2.24	-0.41	-0.43	-0.39	-0.76
Japan	-0.58	-0.58	-0.63	-1.01	-0.58	-1.03	-0.94***	-1.06
Korea	-0.24	-0.33	-0.54	-0.42	-0.34	-0.37	-0.38	-0.38
Netherlands	-	-	-	-	-0.24	-0.25	-0.24	-1.13***
Russia	-0.30	-0.31	-0.53	-0.50	-0.40	-0.67***	-0.41	-0.78***
Spain	-1.07***	-1.29***	-1.29***	-1.08***	-0.19	-0.35	-0.31	-0.61**
Switzerland	-0.89**	-0.87**	-0.89	-0.96*	-0.34	-0.51**	-0.38	-0.54
UK	-0.70***	-0.70***	-0.70***	-0.70***	-0.17	-0.20	-0.35	-0.55
USA	-0.72	-0.74	-0.70	-1.02	-0.46	-0.51	-0.56	-0.96*

Null of no cointegration is tested.
*, **, *** indicate significance at 10%, 5% and 1% levels, respectively. The critical values are taken from Gregory and Hansen, 1996a, p. 109, 1996b, p. 559.

As far as the cointegrated breaking models are concerned, it is worth noting the importance of stability coefficients in the GH approach. Testing one time change in parameters at an unknown date allows the restriction to impose 0.15% equal trimming at both ends of the sample. Table 3 shows the stability coefficients used in export and import demand equations at the bilateral level.

Table 3
Stability Coefficients

	EXPORTS			
	DU	*DUq*	*DUy*	*DT**
Germany	-0.03	-0.96**	-0.07*	0.00
Spain	-0.17	0.37	0.13	0.00
Switzerland	0.08	0.46	0.01	0.00
UK	0.03	-1.17*	-0.01	0.00
	IMPORTS			
Canada	-0.75*	0.79	0.06	0.01
China	-1.33**	1.64	0.50**	-0.01
France	0.00	0.48	0.08	0.01*
Germany	-0.03	-0.96**	-0.07*	0.00
Japan	-0.07	-1.07***	-0.04	0.00
Netherlands	0.16	-1.98***	-0.20***	0.01***
Russia	-107	0.05*	9.00	-0.16*
Spain	-0.62***	1.07***	0.33***	0.03***
Switzerland	0.33	-0.57	-0.05	0.00
USA	-0.06	-0.79*	-0.01	0.00

*, **, *** indicate significance at 10%, 5% and 1% levels, respectively.

DU represents the change in the intercept at the time of the shift. Similarly, *t* represents the time trend before the shift and DT^* represents the time trend at the time of the shift. *DUq* and *DUy* denote the changes in the slope coefficients for the variables such as real exchange rate and real income after the regime shift, respectively. Table 3 shows that stability coefficients are statistically significant in the export demand equation for Germany and the UK, which are the major exporting countries of Turkey. In 2005, 24% of Turkish exports were to Germany and 15% of exports were to the UK. Stability coefficients are statistically significant in import demand equation for all countries other than Switzerland, providing evidence to assume that there is either a level or/and a regime shift in these countries.

6. Data and empirical findings

6.1. Data

The time period for Switzerland is 1983Q1-2007Q4, for Russia 1990Q1-2007Q4 and for all other countries 1982Q2-2007Q4. *RX* is Turkish real exports to country *i*, where nominal export values are deflated from the Turkish export price index (2003=100) obtained from the Turkish Statistical Institute (TURKSTAT). *RM* is Turkish real imports from country *i*, where nominal export values are deflated from the Turkish import price index (2003=100) obtained from TURKSTAT. *Y* is the real GDP in country *i* in dollars, at fixed purchasing power parity based on OECD reference year and seasonally adjusted. Data for GDP in Russia is obtained from Goskomstat and the period 1991-95 is calculated from annual growth rates obtained from the United Nations.

q is the real bilateral exchange rate between Turkey and country *i*. The variable is defined as $(P_{it}^{*}e/P_t)$, where P^* is country *i*'s Consumer Price Index (CPI) (2000=100) is obtained from the OECD for all countries other than China, where it is obtained from the National Bureau of Statistics of China. *e* is the nominal bilateral exchange rate defined as the number of Turkish lira (TL) per number of country *i*'s currency, obtained from the Central Bank of Turkey and *P* is the Turkish CPI[4].

6.2. Empirical findings

This section tries to estimate the long-run and short-run coefficients of trade models between Turkey and its main trading partners. First, the long-run exchange rate elasticities and the income elasticities for both export and import demand functions, respectively. Then, the short-run relations are reported to see whether these variables behave differently.

---

4 Three currencies, the Korean Won, the Chinese Yuan and the Russian Ruble, are calculated from cross rates (via US dollars).

Table 4
Long-run Exchange Rate Elasticities

	C	C/T	C/S	C/S/T
EXPORTS				
Germany	0.45***	0.46***	1.37***	1.32***
Spain	-0.07	-0.14	-0.16	-0.43
Switzerland	0.15	0.12	0.22	-0.15
UK	-0.07	-0.11	0.87	0.92
IMPORTS				
Canada	2.76***	2.71***	2.87***	2.52***
China	-0.56	-0.52	-1.99***	-1.30
France	-0.14	-0.14	-0.16	-0.55**
Germany	0.11	0.07	0.96*	1.32***
Japan	0.22	-0.24	0.45	0.56
Netherlands	-0.22	-0.17	0.69	1.96***
Russia	-0.06***	-0.06***	-0.06***	-0.04**
Spain	0.24**	0.15*	0.95***	-1.12***
Switzerland	-0.17	0.09	0.34	0.39
USA	-0.09	0.14	0.37	0.51**

*, **, *** indicate significance at 10%, 5% and 1% levels, respectively.

The results of the long-run real exchange rate elasticities for the trade functions are reported in Table 4. These results confirm the controversies of different sign and magnitudes in the literature. The real exchange rates are statistically significant only in Germany in the export demand function. Nevertheless, the Turkish export demand to Germany is price elastic for C/S and C/S/T models. The bilateral price elasticity is of the expected sign for Germany. Table 4 also shows that the real exchange rates are statistically significant in the import demand function at least in one of the four models for all countries other than Japan and Switzerland. The import is price elastic in Canada, China, Germany, the Netherlands and Spain. They have expected signs only in China, France, Russia and Spain (in the C/S/T model).

The real exchange rate elasticity of import demand is relatively high in Canada and statistically significant in all GH models. The real exchange rate is elastic showing Turkey's connection to Canadian prices. This is an interesting result, since Turkey has become a significant partner for Canada. Turkish imports of goods from Canada have been increasing steadily over the last decade, such that the value of imports from Canada is well over ten times what it was in the early nineties. Two-way trade between Turkey and Canada was worth about $666 million in 2002 and British Columbia (BC) was the source of over a third of Canadian domestic exports to Turkey. Furthermore, Turkey was the destination for more of BC's goods than either Mexico or Chile, with which Canada has free trade agreements (Schrier, 2003, 2).

Additionally, the real exchange rate elasticity of import demand is relatively low in Russia when compared to the statistically significant real exchange rate elasticity in other countries. One possible explanation is that Turkey's major imports from Russia is energy-related natural resources while the demand for these products are expected to be more inelastic as compared to imports from other countries. Finally, it can be concluded that the ML condition is satisfied only for Germany, Canada, China, Spain, the Netherlands and Spain.

Table 5

Long-run Income Elasticities

	C	C/T	C/S	C/S/T
EXPORTS				
Germany	4.27***	1.72**	4.31***	1.78**
Spain	6.15***	4.43***	6.02***	5.05***
Switzerland	2.05***	0.90	1.99***	2.48***
UK	5.10***	5.29***	5.24***	5.86***
IMPORTS				
Canada	2.68***	2.98***	2.72***	4.22***
China	6.36***	6.73***	5.33***	2.64
France	3.15***	3.19***	3.14***	3.54***
Germany	2.61***	2.80***	2.63***	1.78**
Japan	3.54***	3.17***	2.12***	3.64***
Netherlands	2.51***	2.23***	2.53***	4.17***
Russia	4.51***	1.23*	3.77***	-8.73
Spain	3.63***	0.52	4.54***	2.94***
Switzerland	3.09***	0.24	2.99***	-0.81
USA	1.76***	3.12***	1.90***	3.37***

*, **, *** indicate significance at 10%, 5% and 1% levels, respectively.

Table 3 shows the long-run income elasticities for export and import demand functions, respectively. The results confirm Aydın et al. (2004), Şimşek and Kadılar (2004) and Togan and Berument (2007) with positive sign for both trade demand functions, which is also consistent with the trade demand theories. Income elasticities in both export and import demand functions almost for all models are statistically significant and elastic. Findings in this study show that the income elasticities of both export and import are much greater than one. Additionally, the income elasticities of Turkish exports exceed that of imports for Germany and Spain, which is similar to the findings of Hatemi-J and Irandaust (2005). The study allows including three countries (Germany, Spain and Switzerland) where both export and import estimates are carried out simultaneously. All these countries other than Switzerland gave higher income elasticities of exports than of imports. This shows that Turkish exports are more sensitive to foreign demand and therefore more likely to be responsive to the economic growth in foreign countries.

Table 6

Vector Error Correction Coefficients.

Export Demand Function			
	λ_1	λ_2	λ_3
Germany	-0.04**	0.07***	0.00
Spain	-0.05	-0.14***	0.00
Switzerland	0.02	-0.06***	0.00
UK	0.00	0.07***	0.00
Import Demand Function			
	η_1	η_2	η_3
Canada	-0.02**	0.01***	0.00***
China	0.02	-0.02***	0.00***
France	-0.02	-0.04***	0.01***
Germany	-0.12***	0.12***	-0.03***
Japan	-0.28***	0.16***	-0.02***
Netherlands	-0.12***	0.13***	-0.03***
Russia	-0.03***	-0.02	-0.01***
Spain	-0.01	-0.20***	0.00
Switzerland	0.00	0.02***	0.00***
USA	-0.02	0.06***	-0.02***

*, **, *** indicate significance at 10%, 5% and 1% levels, respectively.
All countries are in lag 1 both for export and import demand functions. Lag selection is according to Schwarz information criterion.

The error correction coefficients, λ and η show the short run behaviours of the trade variables to long run disequilibrium. The error correction coefficients, λ_1 and η_1 , for exports and imports in equations 9 and 10, respectively. λ_1 is negative and significant statistically only in Canada and η_1 is negative and significant statistically in Canada, Germany, Japan, the Netherlands and Russia. The error correction coefficients λ_2 and η_2 show how the relative exchange rates correct long-run disequilibrium. They are significant for all countries for the export demand function and they are significant in all countries other than Russia for the import demand function. Income elasticities, λ_3 and η_3, unlike long-run relationship, have no effect in the adjustment of the long-run disequilibrium.

7. Conclusion

The results of the analysis show that there is a structural change in trade demand function occurred at an unknown date. Structural change is statistically significant, especially in import demand coefficients. This study finds enough evidence to assume a long-run cointegration relationship in export demand function in some of the selected countries, all in Europe. Moreover, a long-run cointegration relationship is found in import demand equations at the bilateral level in most of the selected countries. This shows that the long-run relationship between variables of the trade demand functions exists on a bilateral basis under the assumption of a structural break occurring at an unknown date. Furthermore, the analysis approves there is either a level or/and a regime shift in these countries.

The results of the price elasticities show that, in the long-run, real exchange rates are significant determinants of trade. However, export demand is price elastic only in Germany. The import demand is price elastic in Canada, China, Germany, Spain, the Netherlands and Spain. This emphasises the important role of import demand variables in Turkish trade. The results show that geographical distance is not considered as an important factor in shaping the price elasticity of import demand. Thus, especially the price elasticity of import demand is an important indicator for firms and traders at setting prices, as well as for the Central Bank of Turkey in controlling the current account balance, particularly with those countries.

On the other hand, income is an important determinant of the Turkish trade in the long run. The bilateral income elasticities are much higher than the bilateral price elasticities (real exchange rate elasticities). Both Turkish exports and imports are income elastic. One possible explanation is that income elasticity of trade demand capture all non-price factors excluded from the equation and this can explain the high estimates of income elasticity. Furthermore, Turkish exports are more sensitive to foreign demand, thus more responsive to foreign economic growth.

One robust conclusion is that, in the long run, under the assumption of a structural break, there is a cointegration relationship in trade demand equations on a bilateral basis. But the relations weaken in the short run. A second robust conclusion that has emerged from this study is that the estimated elasticities vary, especially the relative price elasticities, enormously across countries. One possible explanation is that bilateral trade data includes all traded products. However, some traded goods are much easier to substitute than others. For example, commodities are close substitutes, whereas branded goods, cars, and gourmet food are not. The range of traded goods and the share of goods with close substitute may vary from country to country, determining the differences in real exchange rate elasticity.

Nevertheless, this study tests a structural break that occurred at an unknown date in the classical trade demand models for the long run. The study of bilateral trade relations may be carried out with a sector by sector analysis to understand the behaviour of the real exchange rate fluctuations and its effect of trade balance. The analysis may be extended by including multiple structural breaks occurred at an unknown date.

References

ARSLAN, I and VAN WIJNBERGEN, S. (1993), "Export Incentives, Exchange Rate Policy and Export Growth in Turkey", *Review of Economics and Statistics*, 75, 128-133.

AYDIN, M.F., ÇIPLAK, U. and YÜCEL, M.E. (2004), "Export Supply and Import Demand Models for the Turkish Economy", *The Central Bank of the Republic of Turkey Research Department Working Paper*, 04/09.

BAHMANI-OSKOOEE, M. (1986), "Determinants of International Trade Flows: The Case of Developing Countries", *Journal of Development Economics*, 20, 107-123.

BAHMANI-OSKOOEE, M. and BROOKS, T.J. (1999), "Bilateral J-Curve Between US and Her Trading Partners", *Weltwirtschaftliches Archives*, 135, 156-165.

BAHMANI-OSKOOEE M. and GOSWAMI, G.G. (2004), "Exchange Rate Sensitivity of Japan's Bilateral Trade Flows", *Japan and the World Economy*, 16, 1-15.

BAHMANI-OSKOOEE, M, GOSWAMI, G.G. and TALUKDAR, B.K. (2005), "Exchange Rate Sensitivity of the Canadian Bilateral Inpayments and Outpayments", *Economic Modelling*, 22, 745-757.

BAHMANI-OSKOOEE, M. and NIROOMAND, F. (1998), "Long-Run Price Elasticities and the Marshall-Lerner Condition Revisited", *Economics Letters*, 61, 101-109.

CELASUN, M. and RODRIK, D. (1989), "Dept, Adjustment, and Growth: Turkey", Book IV of J. Sachs and S. Collins (eds.), *Developing Country Debt and Economic Performance: Vol.3, Country Studies – Indonesia, Korea, Philippines, Turkey*, Chicago and London, University of Chicago Press.

CHINN, M.D. (2005), "Supply Capacity, Vertical Specialization and Tariff Rates: The Implications for Aggregate US Trade Flow Equations", *National Bureau of Economic Research NBER Working Paper* No W11719.

CUSHMAN, D.O. (1990), "US Bilateral Trade Equations: Forecasts and Structural Stability", *Applied Economics,* 22, 1093-1102.

DOĞANLAR, M. (1998), "Testing for the Structural Break in the Turkish Foreign Trade", *Çukurova Üniversitesi İktisadi İdari Bilimler Fakültesi Dergisi*, 8(1), 333-340.

DICKEY, D.A. and FULLER, W.A. (1979), "Distribution of the Estimators for Autoregressive Time Series with a Unit Root", *Journal of the American Statistical Association*, 74, 427-431.

ENGLE, R.F. and GRANGER, C.W.J. (1987), "Cointegration and Error Correction: Representation, Estimation and Testing", *Econometrica*, 55, 251-276.

GOLDSTEIN, M. and KHAN, M. (1976), "Large Versus Small Price Changes and the Demand for Imports", *IMF Staff Papers*, 23, 200-225.

————(1978), "The Supply and Demand for Exports: A Simultaneous Approach", *Review of Economics and Statistics*, 60, 275-286.

GREGORY, A.W. and HANSEN, B.E. (1996a), Residual-Based Tests for Cointegration in Models With Regime Shifts", *Journal of Econometrics*, 70, 1-26.

————(1996b), "Tests for Cointegration In Models With Regime and Trend Shifts", *Oxford Bulletin of Economics and Statistics*, 58, 555-560.

HATEMI-J, A. and IRANDOUST, M. (2005), "Bilateral Trade Elasticities: Sweden Versus Her Trade Partners", *American Review of Political Economy*, 3, 38-50.

HAYNES, S.E., HUTCHISON, M.M. and MIKESELL, R.F. (1996), "US-Japanese Bilateral Trade and the Yen-Dollar Exchange Rate: An Empirical Analysis", *Southern Economic Journal*, 52, 923-932.

HAYNES, S.E. and STONE, J.A. (1983), "Secular and Cyclical Responses of the US Trade to Income: An Evaluation of Traditional Models", *Review of Economics and Statistics*, 65, 87-95.

HOUTHAKKER, H.S. and MAGEE, S. (1969), “Income and Price Elasticities in the World Trade”, *Review of Economics and Statistics*, 51, 111-125.

IRANDOUST M., EKBLAD, K. and PARMLER, J. (2006), Bilateral Trade Flows and Exchange Rate Sensitivity: Evidence From Likelihood Based Panel Cointegration”, *Economic Systems*, 30, 170-183.

JOHANSEN, S. (1988), “Statistical Analysis of Cointegration Vectors”, *Journal of Economic Dynamics and Control*, 12, 231-254.

KAO, C. and CHIANG, M. (2000), “On the Estimation and Inference of a Cointegrated Regression in Panel Data, Nonstationary Panels, Panel Cointegration, and Dynamic Panels”, in: B. Baltagi, ed., *Advances in Econometrics*, Elsevier, Amsterdam; New York and Tokyo 179–222.

KHAN, M.S. (1974), “Import and Export Demand in Developing Countries”, *IMF Staff Papers,* 21, 678-693.

————(1975), “The Structure and Behaviour of Imports of Venezuela”, *Review of Economics and Statistics*, 57, 221-224.

KREININ, M.E. (1967), “Spurious Regression and Residual-Based Tests for Cointegration in Panel Data”, *Journal of Econometrics*, 90, 510-516.

KWIATKOWSKI, D., PHILLIPS, P.C.B., SCHMIDT, P. and SHIN, Y. (1992), “Testing the null hypothesis of stationarity against the alternative of a unit root: How sure are we that economic time series have a unit root?”, *Journal of Econometrics*, 54, 159-178.

LAL, A.K and LOWINGER, T.C. (2002), “Nominal Effective Exchange Rate and Trade Balance Adjustment in South Asia Countries”, *Journal of Asian Economics*, 13, 371-383.

LIU, L., FAN, K. and SHEK, J. (2007), “Hong Kong’s Trade Patterns and Trade Elasticities”, *Hong Kong Monetary Authority Quarterly Bulletin*, March, 21-31.

MAKIN, A.J. and NARAYAN, P.K. (2008), “Have US External Imbalances Been Determined at Home or Abroad?”, *Economic Modelling*, 25, 520-531.

MARQUEZ, J. (1990), “Bilateral Trade Elasticities”, *The review of Economics and Statistics*, February, 70-77.

NADENICHEK, J. (2000), “The Japan-US Trade Imbalance: A Real Business Cycle Perspective”, *Japan and the World Economy,* 12, 255-271.

NARAYAN, P.K. and NARAYAN, S. (2005), “Estimating Income and Price Elasticities of Imports for Fiji in a Cointegration Framework”, *Economic Modelling*, 22(3), 423-438.

NEYAPTI, B., TAŞKIN, F. and ÜNGÖR, M. (2007), “Has European Customs Union Agreement Really Affected Turkey’s Trade?”, *Applied Economics*, 39, 2121-2132.

ÖZKALE, L. and KARAMAN, F.N. (2006), “Gümrük Birliği’nin Statik Etkileri” *Uluslararası Ekonomi ve Dış Ticaret Politikaları*, 1(1) 117-138.

PERRON, P. (1989), “The Great Crash, the Oil Price Shock and the Unit Root Hypothesis”, *Econometrica*, 57,1361-1401.

————(1997), “Further Evidence on Breaking Trend Functions in Macroeconomic Variables”, *Journal of Econometrics*, 80, 355-385.

PHILLIPS, P.C.B and PERRON, P. (1988), "Testing for a Unit Root in Time Series Regression", *Biometrika*, 75, 335–346.

SCHRIER, D. (2003), “Canada Talks Trade with Turkey”, *BC Stats*, British Columbia Ministry of Management Services, Exports- March 2003.

SCHWERT, G.W. (1989), “Tests for Unit Roots: A Monte Carlo Investigation”, *Journal of Business and Economic statistics*, American Statistical Association, 7, 147-159.

ŞIMŞEK, M. and KADILAR, C. (2004), “Türkiye’nin İhracat Talebi Fonksionunun Sınır Testi Yöntemi ile Eşbütünleşme Analizi”, *Dogus University Journal*, 5, 27-34.

THOMAKOS, D.D. and ULUBAŞOĞLU, M.A. (2002), “The Impact of Trade Liberalisation on Import Demand”, Journal of Economic and Social Research, 4(1), 1-26.

TOGAN, S. and BERUMENT, H. (2007), "The Turkish Current Account, Real Exchange Rate and Sustainability: A Methodological Framework", *The Journal of International Trade and Diplomacy*, 1, 155-192.

UTKULU, U. and SEYMEN, D. (2004), "Trade and Competitiveness Between Turkey and the EU: Time Series Evidence", *Turkish Economic Association Discussion Paper* 2004/8.

UZ, I. KETENCİ, N. (2008), "Panel Analysis of the Monetary Approach to Exchange Rates: Evidence From Ten EU Members and Turkey", *Emerging Markets Review*, 9, 57-69.

WANG, J. and JI, A.G. (2006), "Exchange Rate Sensitivity of China's Bilateral Trade Flows", *BOFIT Discussion Papers*, 19.

WARNER, D. and KREININ, M.E. (1983), "Determinants of International Trade Flows" *Review of Economics and Statistics*, 65, 96-104.

WILSON, J.F and TAKACS, W.E. (1979), "Differential Responses to Price and Exchange Rate Influences in the Foreign Trade of Selected Industrial Countries", *Review of Economics and statistics*, 61, 267-279.

ZIVOT, E. and ANDREWS, D.W.K. (1992), "Further Evidence on the Great Crash, the Oil-Price Shock, and the Unit Root-Hypothesis", *Journal of Business and Economic Statistics*, 10, 251-270.

Özet

Türkiye'nin iki taraflı ticaret esnekliklerindeki yapısal değişimin test edilmesi

Bu çalışma Türkiye'nin temel ticari ortakları ile gerçekleştirdiği çift yönlü dış ticaret esnekliğini Gregory ve Hansen (GH) yaklaşımıyla incelemektedir. Bu nedenle, GH yaklaşımı bilinmeyen bir zamanda yapısal kırık ihtimalini modele dahil ederek uzun dönemde, çift yönlü ihracat ve ithalat talebi fonksiyonlarındaki değişkenler arasında eşbütünleşik ilişki bulmuştur. Ticaret talebi döviz kuru esnekliği ile gelir esnekliğinin 1982-2007 yıllarını kapsayan ampirik sonuçları sunulmuş ve tartışılmıştır.

Anahtar kelimeler: Yapısal değişim, gelir esnekliği, döviz kuru esnekliği, eşbütünleşme.

JEL kodları: F11, F14.