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The Dynamic Effect of Exchange-Rate Volatility on Turkish Exports: Parsimonious Error-Correction Model Approach

Summary: This paper aims to investigate the effect of exchange-rate stability on real export volume in Turkey, using monthly data for the period February 2001 to January 2010. The Johansen multivariate cointegration method and the parsimonious error-correction model are applied to determine long-run and short-run relationships between real export volume and its determinants. In this study, the conditional variance of the GARCH (1, 1) model is taken as a proxy for exchange-rate stability, and generalized impulse-response functions and variance-decomposition analyses are applied to analyze the dynamic effects of variables on real export volume. The empirical findings suggest that exchange-rate stability has a significant positive effect on real export volume, both in the short and the long run.

Key words: Exchange-rate volatility, Export function, Parsimonious error-correction model, Cointegration, Turkey.

JEL: F10, F31.

Stable exchange-rate policies minimizing volatility in the values of national currencies, especially with the difficult global financial conditions at present, have a great impact on national economies. Exchange-rate volatility is emphasized as a disadvantage of flexible exchange-rate regime owing to the nature of daily fluctuations in the nominal exchange-rate that provide reminders of currency risk inherent in flexible regimes (see Emilija Beker 2006). Also, the influence of volatility on world trade may be manifest through its effect on exports; therefore, implementation of stable exchange-rate policy is necessary for providing stability in world trade. Turkey is a country that has made significant recent economic progress, with exports an important part of this progress. According to World Trade Organization (2012), Turkey's share of total world exports in 2000 was 0.41%, which increased to 0.74% in 2011. This moved Turkey from 41st to 32nd position in the list of global exporters. Also, the effect of stable exchange-rate policy on exports has become an important issue for countries with wide current-account deficits. This is because exchange-rate volatility increases financial fragility. Presently, despite taking several measures to combat this problem, the current-account deficit is still an issue in Turkey. The current-account deficit as a percentage of GDP in 2011 and 2012 was 9.7% and 6.1%, respectively.

It is worth noting that exchange-rate volatility affects imports as well as exports. It is expected that exchange-rate volatility negatively impacts imports as it leads to uncertainty about profits gained from international trade, which stems from the lag between the time of trade contact and the payment made in the future time. In this situation, importers may either deflect this uncertainty to the prices dominated in terms of national currency, which would protect them against the possibility of decreasing profits as a result of exchange-rate volatility, or reduce their imports. Although exchange-rate volatility affects imports as mentioned above, we focus on its effect on exports. This is because export-led growth strategies and policies are at the forefront in the medium term programmes in Turkish economy, which are prepared by Ministry of Development for three years.

The aim of this study is to investigate the effect of exchange-rate volatility on Turkish exports in the period of a managed floating exchange-rate regime and therefore, to determine whether the policies based on this strategy and aiming at minimizing the fluctuations in exchange-rate are feasible in Turkish economy. In this study, the effect of exchange-rate stability on exports is investigated by means of cointegration method and parsimonious error-correction model. A generalized impulse-response function and variance-decomposition analyses are also used to determine the dynamic policy effects on exports. This study differs from some of the previous studies investigating the relationship between exchange-rate volatility and Turkish exports in a few ways. First, the parsimonious error-correction model is used in establishing the short-run relationship, and dynamic relationships are investigated by using the impulse-response function and variance-decomposition analyses. Second, the period after the crisis experienced in 2001 in Turkey is analyzed; therefore, the effect of exchange-rate stability on exports is determined by the managed floating exchange-rate regime applied in the post-crisis period. Finally, the Turkish economy in the period discussed evinces a more stable appearance than in previous periods thereby attaining more reliable results in a stable economic environment.

The rest of the paper is organized as follows: in the next section, we present the econometric model used in the analysis. In the second section, we give a brief literature review on exchange-rate volatility and exports relationship. In the third section, we discuss the econometric methodology framework and data source. In the fourth section, we present empirical results. At the end of this paper, we provide some concluding remarks and policy implications.

1. Econometric Model

Many studies analyzing the effect of exchange-rate volatility on exports construct the real export model as a function of variables such as real foreign income, relative price, and exchange-rate volatility. These variables are commonly accepted in the theoretical and empirical literature when constructing the real export model. In this context, the long-run real export demand function, by considering Abdur R. Chowdhury (1993), Augustine C. Arize (1995), Don Bredin, Stilianos Fountas, and Eithne Murphy (2003) and Arize, Thomas Osang, and Daniel J. Slottje (2008), may be expressed as follows:

$$E_t = \beta_0 + \beta_1 w_t + \beta_2 P_t + \beta_3 v_t + \varepsilon_t \qquad (1)$$

where, E_t denotes the natural logarithm of real export volume; w_t is the natural logarithm of real foreign income; P_t is the natural logarithm of relative price; v_t is the natural logarithm of exchange-rate volatility, as a proxy for exchange-rate stability; ε_t is the error term. Economic theory suggests that the increase in real foreign income increases the volume of real exports, and the increase in relative price decreases the volume of real exports by causing domestic goods to be less attractive than foreign goods. Therefore, the expected signs of β_1 and β_2 are positive and negative, respectively. Exchange-rate volatility in the real export demand function may affect real exports both positively and negatively. The models created for the negative effect of exchange-rate volatility on exports are mostly based on risk aversion. Traditional foreign-trade theory suggests under the risk-aversion assumption that exchange-rate volatility will increase uncertainty in the profits expected to be gained from trade, and therefore decreases the trade. As is known, risks encountered while trading is a factor in decreasing trade. Exchange-rate volatility is a significant factor in creating these risks.

In general, trade payments are made on delivery of the product, not in the period when the agreement is signed. Exchange-rate volatility creates uncertainty relating to profit, especially by preventing the prediction of prices and costs, and thereby causes a decrease in international trade and a focus on domestic production (the substitution effect). International traders in developing countries where forward markets are not highly developed are confronted with especially high exchange-rate risk. The development of forward markets, however, decreases the effect of exchange-rate volatility.

2. Literature Review

Bredin, Fountas, and Murphy (2003) suggested that exchange-rate volatility increases the variance of profit, thereby reducing export volume by increasing risk-aversion behavior. According to Chowdhury (1993), an increase in exchange-rate volatility brings additional costs to risk-averse exporters and importers; therefore, domestic trade is preferred. Josef C. Brada and Jose A. Mendez (1988) suggested that trade barriers are created in order to counter the negative effects of unexpected changes in the exchange-rate, thereby decreasing the volume of international trade. Peter Hooper and Steven W. Kohlhagen (1978) also suggested that exchange-rate volatility has a negative effect on trade. Owing to the increase in risk encountered in international trade, traders with a risk-aversion tendency decrease their trade volumes, especially in the absence of hedging facilities, or if existing hedging facilities are costly.

Some economic approaches, however, state that exchange-rate volatility has a positive effect on exports. For example, Paul de Grauwe (1988) suggested that exchange-rate volatility has a positive effect on exports, even in the case of risk-aversion. The exporters with the greatest risk-aversion tendency are likely to ship more exports, since they expect that their incomes will decrease; therefore, trade volume increases. The exporters with a low risk-aversion tendency, however, make fewer efforts to increase exports. Piet Sercu and Cynthia Vanhulle (1992) suggested

that an increase in exchange-risk increases the value of the exporter firm and decreases the exchange-rate at which they leave the market. Udo Broll and Bernhard Eckwert (1999) also asserted that exchange-rate volatility, considered as an option, makes exports more profitable.

Some studies have analyzed the impact of forward exchange markets on the relationship between exchange-rate volatility and trade. Wilfred Ethier (1973) and David P. Baron (1976) suggested that the forward markets' exchange-rate volatility did not affect trade volume. Jean-Marie Viaene and Casper G. de Vries (1992), however, asserted that even in the presence of forward markets, exchange-rate volatility may affect the trade volume indirectly by means of affecting the forward rate. It is worth noting that Agathe Côté (1994), Michael D. McKenzie (1999) and Mohsen Bahmani-Oskooee and Scott W. Hegerty (2007) have provided a comprehensive review about the effect of exchange-rate volatility on exports, indicating different results in theoretical and empirical literature. In summary, the influence of volatility on exports is mixed, from both theoretical and empirical points of view; therefore, exchange-rate volatility has an ambiguous effect on the demand for export goods; that is, the expected sign of β_3 can be negative or positive.

Empirical studies regarding the effect of exchange-rate volatility on exports are carried out by using various econometric methods. Of these methods, the panel-data method is used to determine the relationship between exchange-rate volatility and exports. Christine Sauer and Alok K. Bohara (2001), Gerardo Esquivel and Felibe Larrain B. (2002), Silvana Tenreyro (2007), Joseph P. Byrne, Julia Darby, and Ronald MacDonald (2008) and George Hondroyiannis et al. (2008) can be given as examples of studies using panel-data method. When empirical literature is analyzed, however, it is found that the studies using time-series analysis are weighted. Some of these studies are carried out for more than one country. Examples of these studies are as follows: Arize, Osang, and Slottje (2000) and Arize, John Malindretos, and Krishna M. Kasibhatla (2003) for less developed countries; Arize, Osang, and Slottje (2008) for Latin America countries; Chowdhury (1993) for G-7 countries; Saang J. Baak, Mohammed A. Al-Mahmood, and Souksavanh Vixathep (2007) for East Asian countries. Some of the studies using time-series analysis have also been carried out for a single country. Examples of these studies are: Reza Y. Siregar and Ramkishen S. Rajan (2004) for Indonesia; Pal Boug and Andreas Fagereng (2010) for Norway; Susan Pozo (1992) for Britain; Sijia Zhang and Joseph Buongiorno (2010) for the US; Bahmani-Oskooee and Hegerty (2008) for Japan; Fountas and Bredin (1998) and Bredin, Fountas, and Murphy (2003) for Ireland; Teuku Rahmatsyah, Gulasekaran Rajaguru, and Siregar (2002) for Thailand; Olugbenga A. Onafowora and Oluwole Owoye (2008) for Nigeria.

The Vector Error Correction (VEC) model is mainly used in time-series studies. Chowdhury (1993), Fountas and Bredin (1998), Bredin, Fountas, and Murphy (2003), Baak, Al-Mahmood, and Vixathep (2007), Onafowora and Owoye (2008) and Kosta Josifidis, Jean-Pierre Allegret, and Beker Pucar (2009) are examples of these studies. The ARDL method is also used in some studies like Bahmani-Oskooee and Hegerty (2008), and Florian Verheyen (2012). When the results of empirical studies are evaluated in a general sense, it is observed that some studies find the nega-

tive effect of exchange-rate volatility on exports for different countries. Jerry G. Thursby and Marie C. Thursby (1987), Chowdhury (1993), Kiheung Kim and WooRhee Lee (1996), Arize, Osang, and Slottje (2000), W. L. Chou (2000), Bahmani-Oskooee (2002), Murat Doğanlar (2002), Zhang and Buongiorno (2010) and Verheyen (2012) are examples of such studies. Hooper and Kohlhagen (1978), Bahmani-Oskooee and Sayeed Payesteh (1993), Arize (1999) and Eleanor Doyle (2001), however, concluded only an insignificant or positive correlation between exchange-rate volatility and exports.

In the empirical literature concerning developing and emerging countries, there are a large number of studies investigating the relationship between exchange-rate volatility and exports. The econometric methods and findings obtained from some of these studies are represented as follows: Arize, Osang, and Slottje (2000) investigated the impact of real exchange-rate volatility on the export flows of 13 less developed countries (LDCs) using quarterly data over the period 1973-96, by using the error-correction model. The majority results show that increases in the volatility of the real effective exchange-rate have a significant negative effect on export demand in both the short and long run in each of the 13 LDCs. The other study of LDCs carried out by Arize, Malindretos, and Kasibhatla (2003) covered quarterly data for the period 1973-1998, and empirical results of cointegration tests and error-correction models indicated that increases in exchange-rate volatility have a significant negative effect on export demand in both the short and long run in most of the countries studied.

Josifidis, Allegret, and Beker Pucar (2011) investigated the features of a managed floating exchange-rate regime within the monetary framework of inflation targeting in case of Serbia and other selected transition economies (Poland, the Czech Republic and Slovakia) by using VEC and Vector Autoregression (VAR) models for different periods. They found that (in)direct management of exchange-rate fluctuations was relatively highest in Serbia, compared with other cases, which was connected with the highest exchange-rate pass-through and financial euroization. Baak, Al-Mahmood, and Vixathep (2007) investigated the impact of exchange-rate volatility on exports from four East Asian countries (Hong Kong, South Korea, Singapore and Thailand) to the US and Japan using annual data from 1981 to 2004. According to cointegration tests and error-correction models, except in the case of Hong Kong's exports to Japan, exchange-rate volatility has a negative effect on exports either in the short or the long run. Stephen Hall et al. (2010) used a generalized method of moments estimation and time-varying-coefficient estimation for the emerging-market economies and developing countries, using data for the period 1980:Q1-2005:Q4 and 1980:Q1-2006:Q4. According to the empirical findings, for the developing countries, exchange-rate volatility negatively affects exports whereas for the emerging-market economies, the findings do not provide support for the hypothesis that exchange-rate volatility has a significant negative effect on exports. Serge Rey (2006) investigated the impact of nominal and real effective exchange-rate volatility on the exports of six Middle Eastern and North African countries, including Turkey, to fifteen member countries of the European Union (EU), for the period 1970:Q1-2002:Q4. According to the cointegration results, there is a significant relationship, negative for Algeria,

Egypt, Tunisia and Turkey, positive for Israel and Morocco, between MENA exports and exchange-rate volatility. Onafowora and Owoye (2008) investigated the relationship between exchange-rate volatility and Nigeria's exports using quarterly data from January 1980 to April 2001. The estimates of the cointegration test and error-correction model indicated that real exchange-rate volatility had a negative effect on exports, both in the short and long run. Siregar and Rajan (2004) investigated the relationship between exchange-rate volatility and Indonesia's exports for the period 1980:Q2-1997:Q2, and the cointegration test results indicated that exchange-rate volatility had a negative effect on exports. Arize (1999) and Arize, Osang, and Slottje (2008) investigated the impact of real exchange-rate volatility on export flow on a quarterly basis over the period 1973-2004 for eight Latin American countries and over the period 1973:Q2-1997:Q1 for Singapore, respectively. The results of the cointegration tests and error-correction models show that increases in the volatility of the real effective exchange-rate have a significant negative effect upon export demand in both the short and the long run in each of the eight Latin American countries, but have a statistically significant positive effect on real exports for Singapore. As regards the studies investigating the effect of exchange-rate volatility on exports by using Turkish data, the explanations of the methods and empirical results of these studies are reported in the Appendix 1.

3. Methodology Framework and Data Source

In this study, in order to determine whether the variables in Equation (1) have a unit-root, Augmented Dickey-Fuller test (ADF) (David A. Dickey and Wayne A. Fuller 1981), Phillips-Perron test (PP) (Peter C. B. Phillips and Pierre Perron 1988) and KPSS test (Denis Kwiatkowski et al. 1992) are employed. If the variables are stationary in their first differences, the long-run relationship among the variables is tested by using the cointegration method, developed by Soren Johansen (1988, 1991) and Johansen and Katarina Juselius (1990, 1992). For testing cointegration vectors, trace and maximum statistical eigenvalues are established. In addition to cointegration analysis, the parsimonious model, obtained from the general error-correction model and demonstrated in Equation (2), is estimated to determine the short-run effect of exchange-rate volatility on real export volume. For establishing the parsimonious error-correction model, David F. Hendry's (1974, 1977) "general to specific" approach is applied. In this approach, first, the error-correction model is estimated, and then, statistically insignificant variables are excluded from the general error-correction model.

$$\Delta E_t = \alpha \sum_{i=0}^{n1} \beta_i \Delta E_{t-i-1} + \sum_{i=0}^{n2} \varphi_i \Delta w_{t-i} + \sum_{i=0}^{n3} \delta_i \Delta P_{t-i} + \sum_{i=0}^{n4} \psi_i \Delta v_{t-i} + \lambda_i EC_{t-1} + \mu_t \qquad (2)$$

where, Δ is the first difference operator; n is the number of lags involved in the estimation; EC_{t-1} is the error-correction term obtained from the cointegration equation; u_t is the disturbance term. E_t, w_t, P_t, and v_t are real export volume, foreign real income, relative price, and exchange-rate volatility, respectively. All data are expressed in

natural logarithmic form. If the variables are not cointegrated, EC_{t-1} is excluded from Equation (2) and the equation is estimated without EC_{t-1}. In addition to these analyses, generalized impulse-response functions (GIRFs) and variance-decomposition analysis are carried out in order to determine the dynamic effect of variables on real export volume; therefore, the effect of one standard deviation shock occurring in exchange-rate volatility, foreign income, and relative price on the real export volume may be determined.

In this study, we selected the period between February 2001 and January 2010. The reason for selecting this period is that it is covered by a managed floating exchange-rate regime, adopted as an exchange-rate regime after the crawling-peg regime was abandoned in February 2001. Export data are expressed in real terms by using the US consumer price index and adjusting seasonally. The reason for using the US consumer price index to make export data real is that it was obtained in dollars. As an indicator of foreign income, the average of the seasonally adjusted industrial production index of 26 EU countries is taken. The reason for taking the incomes of EU countries as the foreign-income indicator is that they have an important place among the countries to which Turkey exports; for instance, the share of Turkish exports going to EU countries in 2011 was 46.2%.

The relative price variable is represented by the international terms of trade, calculated by dividing the export-price index by the import-price index. As regards the sources of data, the US consumer price index used for expressing export series in real terms was obtained from the International Monetary Fund (IMF) database; industrial production indices of EU countries were obtained from the EUROSTAT database; and other variables were obtained from the electronic data delivery system of the Central Bank of the Republic of Turkey (CBRT). The variables used in the models and their explanations are presented in the Appendix 2.

Exchange-rate volatility is not a directly observable variable. However, there are several methods for measuring it. These methods are the moving sample standard deviation of growth rate of exchange-rates (see Chowdhury 1993; Arize, Osang, and Slottje 2000; Bahmani-Oskooee 2002); percentage change in exchange-rates (see Martin J. Bailey, Tavlas, and Ulan 1987); residuals from the ARIMA model (see Ahmed A. A. Asseery and David A. Peel 1991); and GARCH models (see Pozo 1992; Kenneth F. Kroner and William D. Lastrapes 1993; Tony Caporale and Khosrow Doroodian 1994; Jaewoo Lee 1999; Helmut Herwartz 2003; Chongcheul Cheong, Tesfa Mehari, and Leighton V. Williams 2005; Taufiq Choudhry 2005, 2008; Zhang and Buongiorno 2010).

In this study, nominal exchange-rate volatility of the Turkish Lira (*TL*) against the US dollar, *TL/US* dollar parity, is estimated by means of the GARCH model. After applying the GARCH (p, q) model to a nominal exchange-rate series, the resulting conditional variance estimate is used to determine exchange-rate volatility. For this purpose, first, ADF, PP and KPSS unit-root tests are applied for the natural logarithm of the nominal exchange-rate series; and if the unit-root tests show that the exchange-rate series were I(0) in log-level, then the GARCH model is estimated for the log-level. When estimating the GARCH (p, q) model, the most suitable AR (m) model for the natural logarithm of nominal exchange-rate is determined by consider-

ing Akaike's Information Criterion (AIC) and Ljung-Box Q statistics for the absence of serial correlation in residuals. Considering the results, we determine the most suitable model as AR (2); and after estimating the AR (2) model, we apply the ARCH-LM test for investigating the absence of heteroscedasticity. According to ARCH-LM test results, the ARCH effect is found in the residuals; therefore, the GARCH (1, 1) model is applied, and the square roots of conditional variance are estimated in order to obtain the series of exchange-rate volatility.

It is worth noting that a large number of models have been developed to measure volatility in the empirical literature as well as in the GARCH models. Regarding the suitability of the GARCH (1, 1) model for measuring volatility, Robert F. Engle (2001) states that the ARCH and GARCH models are used for the purpose of providing volatility measurement. Such models can be used in financial decisions concerning risk analysis, portfolio selection, and derivative pricing. Engle (2001) also suggests that the GARCH (1, 1) model is the simplest and most robust of the family of volatility models and therefore, has proved sufficient for most financial-market data. According to Vivek Bhargava and Davinder K. Malhotra (2007), numerous previous studies have shown that the conditional variance of the GARCH (1, 1) model is the appropriate volatility measure for currencies. Kenneth D. West, Hali J. Edison, and Dongchul Cho (1993) also demonstrated that the GARCH (1, 1) model displays a better performance than ARCH models. Additionally, Peter R. Hansen and Asger Lunde (2005) compared different ARCH-type models in terms of their ability to describe the conditional variance for both exchange-rate and IBM stock returns. According to the empirical findings, there is no evidence that the GARCH (1, 1) model is outperformed by other models for exchange-rate data and there is evidence that the GARCH (1, 1) model is inferior to other models for the stock-returns data. The GARCH specification takes the form:

$$y = \beta_0 + \beta_1 y_{t-1} + \beta_2 y_{t-2} + e_t \tag{3}$$

where, y_t is the natural logarithm of the *TL/US* dollar parity for period t and

$$\sigma_t^2 = \alpha + \beta e_{t-1}^2 + \gamma \sigma_{t-1}^2 + \mu_t \tag{4}$$

where, u_t is a white-noise term. The conditional variance equation (Equation 4) is a function of three terms: (1) the mean (α); (2) the news about volatility from the previous period effecting exchange-rate volatility, measured as the lag of squared residual from the mean equation, e_{t-1}^2 (the ARCH term); (3) the last period's forecast-error variance σ_{t-1}^2, the GARCH term, lagged conditional variance, measuring the effect of forecast variance from previous periods on the current conditional variance. Conditional variance estimated from Equation (4) is the measurement of exchange-rate volatility applied in the estimation of Equation 1.

For a GARCH (1, 1) process, $\alpha > 0$ and both β and γ should be non-negative, $\beta \geq 0$ and $\gamma \geq 0$, and $\beta + \gamma < 1$. The sum of β and γ measures the persistence of volatility shocks. The sum of β and γ measures the persistence of volatility shocks. If the sum of β and γ is less than one, the shock reduces over time; that is, the unconditional va-

riance is finite, and stationarity is ensured by not having a unit-root, as shown by Tim Bollerslev (1986). If the sum of the parameters is equal to one, however, the shock will affect volatility for an uncertain time. Finally, closing the sum of these variables to one implies that the shocks will affect volatility for some time in the future. The GARCH (1, 1) method estimates are represented in Equation (5); standard errors are in parentheses. According to the ARCH-LM (1) test, applied to the estimated GARCH (1, 1) residuals, there is no evidence of conditional heteroscedasticity.

$$\sigma_t^2 = 0.0005 + 0.89e_{t-1}^2 + 0.006\,\sigma_{t-1}^2 + \mu_t.$$
$$(0.0001)\quad(0.243)\qquad(0.084)$$

(5)

The results of the variance equation indicate that the restrictions put on the coefficients stated above are satisfied. Satisfying the condition $\beta + \gamma < 1$ implies finite conditional variance, with the shocks reducing over time. The coeffecient of α is small and significant, and although the estimation of lagged squared residual (e_{t-1}^2) is significant, lagged variance (σ_{t-1}^2) is insignificant. According to Choudhry (2005), the half-life of a shock to volatility is measured by 1 - [log 2 / log ($\beta + \gamma$)]. Half-life measures the period of time (number of months) over which a shock to volatility reduces to half its original size. Using this formula, the half-life of the shock is calculated as 7.3 months.

A graphical representation of the volatility of the *TL/US* dollar parity is presented in Figure 2 (see Appendix 3). From this, we conclude that the parity is highly volatile since the adoption of a floating exchange-rate regime in early 2001. For instance, the exchange-rate exhibited high volatility at the beginning of 2001, in the middle of 2004, in June and July 2006, and in November 2008. The reason behind high volatility in the early months of 2001 was the rapid increase and high volatility in *TL/US* dollar parity after the floating exchange-rate regime adopted on 22nd February 2001. During 2001, *TL* depreciated by 115.3% and 107.1% against the US dollar and the euro, respectively. In May 2004, owing to the widening current-account deficit and the increasing interest rate in the US, *TL* depreciated by 11% against the exchange basket, consisting of US$1 and €0.77. At the beginning of 2006, the increase in interest rate owing to inflationary environment in developed countries and rising risk-aversion tendency in global markets had a negative effect on net international capital inflows. In mid-2006, the volatility in *TL/US* dollar parity moved to a high level, and *TL* depreciated by more than 20% against the US dollar. Finally, the deepening of the global financial crisis and the subsequently increasing risk prime in 2008 also increased the volatility of *TL/US* dollar parity.

Summary statistics and the correlations of all variables used in the study are presented in Appendix 4. The symmetry of data is measured by skewness, and whether the data are peaked or flat relative to a normal distribution is measured by kurtosis. Skewness value 0 and kurtosis value 3 indicate that the variables are normally distributed. The skewness values for variables used in the study vary between -0.56 and 1.36. The distribution of *E* and *P* are negatively skewed (that is, skewed to the left) and that of *w* and *v* are positively skewed (that is, skewed to the right). The kurtosis values for variables vary between 1.89 and 4.58. The kurtosis value of *v* is greater than that of other variables, indicating that *v* is more leptokurtic or fat-tailed

than the other variables. The relatively low kurtosis values of the other variables indicate that the distributions are flat, near the mean for these variables.

According to the Jarque-Bera test - a formal test of normality - results, the null hypothesis of normality is rejected for all variables. Given the results of all tests, it is obvious that all the variables used in the study are non-normally distributed. Appendix 4 also presents the correlations between variables. It is clear that the correlations between independent variables are low in the studying period. E, however, is highly positively related to w (0.84), but negatively related to P (-0.58) and v (-0.38). Additionally, it is concluded that the international terms-of-trade variable (P) is less volatile than other variables (with a standard deviation of 0.03).

Prior to estimation, we used ADF and PP tests to determine the unit-root properties of the series. The KPSS test was also performed because of the small sample size and the fact that ADF and PP tests are known to have small power. The KPSS tests for a stationary null hypothesis contrary to ADF that tests for the null of a unit-root. According to the unit-root test results, represented in Appendix 5, E and w are not stationary in levels, but stationary in first differences in all tests; that is, E and w are I(1). P and v are, however, stationary in level, I(0), according to the ADF and PP tests and stationary in their first differences, l(1), according to the KPSS test.

4. Empirical Results

4.1 Cointegration Tests

After obtaining the unit-root test results, we considered the KPSS test because of the shorter period and absence of a structural break, concluding that all variables are stationary in their first difference. Therefore, for determining the existence of a long-run relationship between the variables, the cointegration test, as proposed by Johansen (1988, 1991) and Johansen and Juselius (1990, 1992), is applied. The lag length is determined by the Akaike's Information Criterion and Schwartz Bayesian Criterion. Maximal Eigenvalue and Trace tests are given in Table 1, Panel A. The results of these tests show that the null hypothesis of no cointegration is rejected, but that of at most one cointegration vector is not rejected. This indicates the presence of one cointegrating relationship and thus shows the long-run relationship between real exports, foreign income, relative price, and exchange-rate volatility. Table 1, Panel B, shows the long-run elasticities obtained from the normalized cointegration equation.

According to the results, a 1% increase in exchange-rate volatility led to a 0.63% decrease in real export volume in the long-run, and the effect is statistically significant at the 1% level, in accordance with expectations. A 1% increase in foreign income led to a 2.63% increase in real export volume in the long-run, and the effect is statistically significant at the 10% level, in accordance with expectations. As for the relative price, it has a positive but insignificant effect on real export volume, as opposed to expectations.

Table 1   Results of Johansen Cointegration Tests

Panel A					
Hypothesis	**Trace**	**5% critical value**	**Hypothesis**	**Max-eigen**	**5% critical value**
H_0: r = 0	50.10**	47.85	H_0: r = 0	28.54**	27.58
H_0: r ≤ 1	21.55	29.79	H_0: r ≤ 1	14.83	21.13
H_0: r ≤ 2	6.72	15.49	H_0: r ≤ 2	4.45	14.26
H_0: r ≤ 3	2.26	3.84	H_0: r ≤ 3	2.26	3.84

Panel B
The normalized cointegration equation

$$E = 0.21P + 2.30w - 0.63v$$
$$(3.82) \quad (1.35) \quad (0.10)$$

Notes: In the cointegration equation, the model with a constant is used; r: number of cointegration vectors. Optimum lag length is found to be 3 on the basis of Akaike's Information Criterion and Schwartz Bayesian Criterion. ** expresses that the null hypothesis suggesting there is no cointegration is rejected at the significance level of 5%. Values within the parenthesis under the estimated coefficients express the standard errors. The critical values for the Johansen tests are based on, James. G. MacKinnon, Alfred A. Haug, and Leo Michelis (1999).

Source: Authors' estimations.

4.2 Parsimonious Error-Correction Model

Although the cointegration test results indicate a long-run relationship among the variables, they do not provide any such indication concerning the short-run dynamics. The theory developed by Engle and Clive W. J. Granger (1987) states that an error-correction model may be established if there is a cointegration relationship among the variables. The error-correction model is created by adding one period-lagged value of the error terms obtained from the cointegration equation and the first differences of the variables into the regression; therefore, the short-run real export function is analyzed. First, the error-correction model is established as 12 lagged, and then Hendry's (1974, 1977) general to specific approach is applied; that is, the variables that are not statistically significant are omitted from the general model, and the parsimonious error-correction model is therefore constructed. The coefficients, standard errors, and *t*-statistics in the parsimonious error-correction model are provided in Table 2.

According to the results from the parsimonious error-correction model, in the short-run, in accordance with the expectations, a 1% increase in exchange-rate volatility led to a 0.017% decrease in real export volume; a rise in foreign income of 1% led to 1.533% and 1.700% increases in the real export volume for the first and the second month, respectively. These effects are also statistically significant, at least at the 5% level. For the relative price, however, the effect of exchange-rate volatility has a negative but insignificant effect on real export volume in the short-run, contrary to expectations. To sum up, exchange-rate volatility affects real export volume negatively in the short-run, as well as in the long-run. Foreign income has positive effect on real export volume in the long and short-run. The relative price, however, has no statistically significant effect on real export volume in either the long or the short-run. It is also observed that the error-correction term (EC_{t-1}) is negative and statistically significant, confirming the long-run relationship between real export volume,

foreign income, relative price, and exchange-rate volatility. Additionally, the size of the coefficient of an error-correction term shows that 3.4% adjustment of real exports toward long-run equilibrium occurs per month.

Table 2   The Results from the Parsimonious Error-Correction Model

Variables	Coefficient	Std. error	*t*-statistics
ΔE_{t-1}	-0.902***	0.086	-10.478
ΔE_{t-2}	-0.500***	0.083	-5.973
ΔP_t	-0.311	0.406	-0.764
Δw_t	1.533**	0.699	2.191
Δw_{t-1}	1.700**	0.751	2.262
Δv_t	-0.017**	0.008	-2.020
EC_{t-1}	-0.034**	0.016	-2.095
Diagnostic tests			
Adj. R^2	0.55		
Jarque- Bera	0.54		
Breusch-Godfrey-LM	0.46		
ARCH-LM	0.34		
Ramsey Reset	0.63		

Notes: The asterisks (***) and (**) indicate the rejection of the null hypothesis at the 1% and 5% significance levels, respectively. B-G, J-B and ARCH-LM nulls are no serial correlation, normality and no ARCH up to the selected lag, respectively. Lag lengths are selected by Akaike's Information Criterion and Schwartz Criterion and are in parentheses. White test is with cross terms and null is no heteroscedasticity. Ramsey RESET test null is the one with no specification errors and is conducted for one fitted term using Likelihood Ratio.

Source: Authors' estimations.

The diagnostic tests used to determine the robustness of the parsimonious error-correction model are shown at the bottom of Table 2. These tests are: the Jaque-Berra test for normality of residuals; the Breusch-Godfrey LM test of residual serial correlation; the ARCH-LM test for heteroscedasticity; and the Ramsey Reset test for specification error. The *p* values of all tests are higher than 0.05; therefore, the Jargue-Bera, ARCH-LM, Breusch-Godfrey LM serial correlation, and Ramsey Reset tests show normally distributed error terms, no autoregressive conditional heteroscedasticity, no autocorrelation problems, and no specification error in the model, respectively. Moreover, the adjusted R^2 of the model is 0.55, which is a fair value. Taking into account the possibility that the crisis in the global economy in 2008 caused a structural break in the Turkish real export function, the CUSUM and CUSUM-SQ statistics of the estimated model are calculated. CUSUM and CUSUM-SQ statistics are within the 95% confidence bands, implying that there is no structural break in the analyzed period. These tests can be obtained from the authors upon request. When the diagnostic tests are evaluated together, there is no sign of model misspecification or parameter instability, and the tests therefore point to the robustness of the model.

4.3 Generalized Impulse-Response Functions and Variance-Decomposition Analysis

In addition to analyzing both long and short-run relationships by using the cointegration method and parsimonious error-correction model, generalized impulse response functions (GIRFs), as proposed by Gary Koop, M. Hashem Pesaran, and Simon M. Potter (1996) and Pesaran and Shin (1998), are used to determine the dynamic interaction of real export volume and its determinants. Impulse responses present the impact of one standard deviation shock or innovation of one variable on the current and future values of another variable. Orthogonalized impulse-response function analysis, introduced by Cristopher A. Sims (1980), depends on the order of the variables in Choleski decomposition, leading to the emergence of different impulse-response results. GIRF analysis, however, is not sensitive to the order of variables and therefore, yields unique impulse-response functions that are unvarying whatever the order of the variables. Optimum lag length is found to be 3 for the ECM on the basis of Akaike's Information Criterion and Schwartz Bayesian Criterion. The responses of real exports to one standard deviation shock in other variables over a 10-month period are presented in Figure 1. From Figure 1, it can be concluded that exchange-rate volatility has persistent negative effect on real exports, the relative price variable has an unpredictable effect, and foreign income increasingly has a positive effect. GIRFs results also support those obtained from the short and long run real export models.
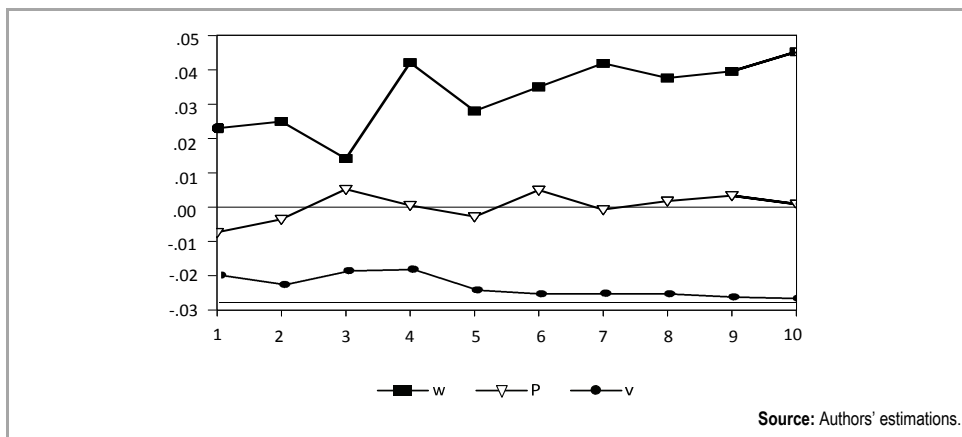

Source: Authors' estimations.

Figure 1 Responses of Real Exports to One Standard Deviation Shock in Other Variables

The variance-decomposition of the real export function shows the relative importance of shocks in explaining the variation in the real export variable. The decomposition of forecast-error variance for a 10-month forecasting horizon for the real export function is presented in Table 3. The order used for the Choleski decomposition is as presented in Table 3. According to the results, at the end of 10 months, foreign income is the most effective variable on real exports, and the shock to exchange-rate volatility has a small effect on real exports in the first few periods, but a

significant one in the later periods. Although exchange-rate volatility explains 6.85% of real exports in the first period, it reaches 18.62% in the tenth period.

Table 3 Variance-Decomposition Analysis Results

Period	v	w	P	E
1	6.85	4.81	0.37	87.95
2	13.75	8.82	0.32	77.09
3	15.30	7.85	0.90	75.94
4	13.54	18.04	0.83	67.57
5	16.01	18.47	0.72	64.78
6	17.41	20.00	1.10	61.46
7	17.28	22.18	0.96	59.55
8	18.31	23.76	1.02	56.88
9	18.52	24.41	1.12	55.93
10	18.62	26.12	1.09	54.15

Source: Authors' estimations.

5. Conclusions and Policy Implications

This paper aims to investigate the effect of exchange-rate stability on real exports in the short and long-run by using the data for the period February 2001 to January 2010. For this purpose, a real export model consisting of foreign income, relative price, and exchange-rate volatility -as a proxy for exchange-rate stability- are used. To determine the long-run relationship among the main variables affecting real export volume, a Johansen cointegration test is carried out. For determining the short-run relationship among the variables, the parsimonious error-correction model is also estimated. The results of the cointegration test indicate that, in the long-run, exchange-rate volatility and foreign income have negative and positive significant effect on real export volume, respectively.

The results of the parsimonious error-correction model created for determining the short-run effects support the long-run results. In this context, exchange-rate volatility and foreign income have negative and positive effect on real export volume in the short-run, respectively. As regards the magnitude of the effects, the coefficient of the foreign-income variable is higher than that of exchange-rate volatility, both in the short and long-run. Relative price, however, has an insignificant effect on real export volume both in the long and short-run. Additionally, GIRFs and variance-decomposition analyses are carried out for investigating the dynamic effects of variables on real export volume. The results indicate that exchange-rate volatility has a persistent negative effect on real export volume, and that the shock to exchange-rate volatility has a substantial effect in the later periods, supporting the results obtained from the parsimonious error-correction model.

These empirical results have important policy implications to assist policymakers in increasing exports from Turkey. Exchange-rate volatility and the national income of the countries with a crucial share in Turkish foreign trade have a significant effect on Turkish real exports. It is also observed that foreign income has a greater effect on real exports than exchange-rate volatility. These findings show the sensitiveness of Turkish real exports to the economic crises faced by European coun-

tries. At this point, policymakers should diversify exports; then the negative effects of the fluctuations in foreign income can be minimized. The empirical results also suggest that measures toward decreasing exchange-rate fluctuations should be taken to increase and stabilize real exports when implementing economic policies, and that policymakers should implement the policies that minimize the exchange-rate fluctuations.

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Appendix 1

Table 4   Summary of Previous Studies for Turkey

Study	Period	The method of volatility measurement	Methodology	Results
İlhan Öztürk and Ali Acaravcı (2002)	1989:01 2002:08	Standard deviation of the percentage change in the real exchange-rate.	Cointegration.	Negative effect of exchange-rate volatility on exports in the long-run.
Pınar Özbay (1999)	1998:Q2 1997:Q2	GARCH method.	Ordinary least squares method.	Negative effect of exchange-rate volatility on exports.
Nezir Köse, Ahmet Ay, and Nurgül Topallı (2008)	1995:01 2008:06	Standard deviation of the percentage change in the real exchange-rate, GARCH method.	Cointegration, error-correction model.	Negative effect of exchange-rate volatility on exports both in the short and long run.
Hasan Vergil (2002)	1990:01 2000:12	Standard deviation of the percentage change in the real exchange-rate.	Cointegration, error-correction model.	Negative effect of exchange-rate volatility to Germany, France, and the US in the long-run. Negative effect of exchange-rate volatility on exports to Germany in the short-run.
Serpil Türkyılmaz, Mustafa Özer, and Erol Kutlu (2007)	1999:01 2007:01	AR (1) - TGARCH (1, 1) method.	Granger causality test of the standard VAR model, impulse-response functions, variance-decomposition analysis.	Unidirectional causality from exports and imports to nominal exchange-rate volatility. The results of the impulse-response function and variance-decomposition analyses also support the causality test results.
Selim A. Hatırlı and Kübra Önder (2010)	1998:2008 (Monthly data)	GARCH method.	Cointegration, Granger causality test of the error-correction model.	Bidirectional causal relationship between exchange-rate volatility and export of textile and apparel industry.
Cumhur Büğük et al. (2001)	1982:Q1 1998:Q4 1986:Q1 1998:Q4	Standard deviation of the percentage change in the real exchange-rate.	Cointegration, error-correction model.	Except a few countries, exchange-rate variability does not have significant impact on Turkish exports of dried figs, grapes and tobacco.
Baki Demirel and Cumhur Erdem (2004)	1990:Q1 1998:Q4	EGARCH method.	Error-correction Model.	Negative effect of exchange-rate volatility on exports in general.

Appendix 2

Table 5   Data Sources and Descriptions

Variable	Definition	Description	Source
E	Export volume	Seasonally adjusted real export volume (using the US consumer price index and expressed in logatithmic form).	IMF CBRT
w	Foreign income	Seasonally adjusted average industrial production index of 26 EU countries (expressed in natural logarithmic form).	EUROSTAT
P	Relative price	Export-price index / Import-price index (expressed in natural logarithmic form).	CBRT
v	Exchange-rate volatility	Volatility of *TL/US* dollar parity measured using GARCH (1, 1) model (expressed in natural logarithmic form).	CBRT

Appendix 3

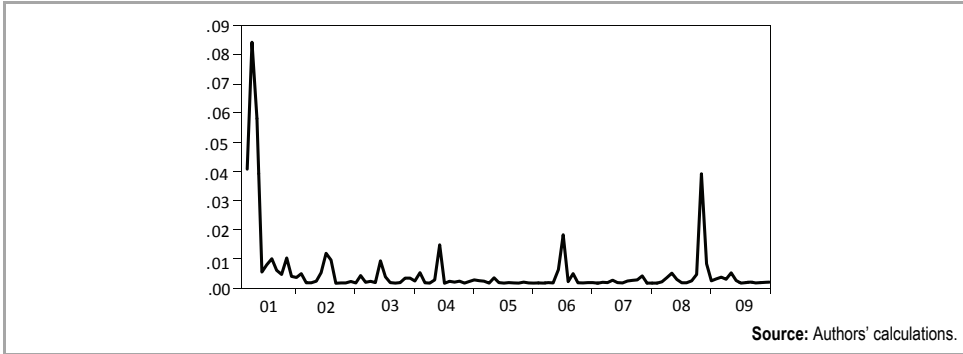**Source:** Authors' calculations.

Figure 2 Volatility of TL/US Dollar Parity, 2001:Q2-2010:Q1

Appendix 4

Table 6 Descriptive Statistics and Correlation Matrix

	E	w	P	v
Mean	4.04	4.59	4.58	-6.52
Median	4.13	4.58	4.58	-6.85
Maximum	4.69	4.73	4.65	-2.48
Minimum	3.30	4.48	4.49	-7.70
SD.	0.38	0.07	0.03	1.15
Skewness	-0.36	0.37	-0.56	1.36
Kurtosis	2.00	1.89	3.25	4.58
Jarque-Bera	6.85**	8.08**	5.97**	44.46***
Observations	108.00	108.00	108.00	107.00
Correlation matrix				
E	1.00			
w	0.84	1.00		
P	-0.58	-0.56	1.00	
v	-0.38	-0.33	0.06	1.00

Notes: *E:* real export volume; *w:* the average industrial production index of 26 EU countries; *P:* international terms of trade; *v:* exchange-rate volatility, measured by the square roots of conditional variance by estimating GARCH (1, 1) model. All data are expressed in natural logarithmic form. SD = standard deviation; JB represents the Jarque-Bera normal-distribution-test statistic. **and *** denote significance at the 5% and 1% levels, respectively.

Source: Authors' calculations.

Appendix 5

Table 7 Estimation of Unit-Root Tests

Variables	ADF		PP		KPSS	
E	-0.49(2)	η_T	-3.27 (5)	η_T	0.22(8)***	η_T
ΔE	13.89(1)***	η_μ	-21.49 (10)***	η_μ	0.22(15)	η_μ
W	-1.77(3)	η_T	-0.78(6)	η_T	0.19(8)**	η_T
Δw	-3.22(2)**	η_μ	-9.48(6)***	η_μ	0.24(7)	η_μ
P	-3.74(0)**	η_T	-3.98(1)**	η_T	0.72(8)**	η_μ
ΔP	-10.12(1)***	η_μ	-9.89(9)***	η_μ	0.05(11)	η_μ
V	-6.04(0)***	η_T	-5.74(5)***	η_T	0.24(5)***	η_T
Δv	-8.65(4)***	η_μ	-16.92(9)***	η_μ	0.21(15)	η_μ

Notes: PP: Phillips-Perron; ADF: Augmented Dickey-Fuller; KPSS: Kwiatkowski, Phillips, Schmidt, and Shin tests. ** and *** means that in ADF and PP tests, the null hypothesis suggesting the series have unit-roots are rejected at the levels of 5% and 1%, respectively; and in KPSS test, the null hypothesis suggesting that the series do not have unit-roots are rejected at the levels of 1% and 5%, respectively. In applying the ADF test, lag lengths are determined on the basis of Akaike Information Criterion, the Schwartz Bayesian Criterion and the Hannan-Quinn Criterion, and shown in parenthesis. The bandwidth in KPSS and PP tests are determined by using the Newey-West method and are shown in parenthesis. η_T expresses a trend model, and η_μ xpresses a non-trend model. In non-trend models; the critical values for ADF and PP tests are -3.49 and -2.88 for 1% and 5%, respectively, and the critical values for the KPSS test are 0.73 and 0.46 for 1% and 5%, respectively. In trend models; the critical values for ADF and PP tests are -4.04 and -3.45 for 1% and 5%, respectively, and the critical values for the KPSS test are 0.21 and 0.14 for 1% and 5%, respectively. The ADF, PP, and KPSS test equations are well known and are therefore omitted to conserve space.

Source: Authors' estimations.

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