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Exchange Rate Volatility in Turkey and Its Effect on Trade Flows

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Abstract. This paper empirically investigates the impact of real exchange rate volatility on the export flows of Turkey to the United States and its three major trading partners in the European Union for the period 1990:1-2000:12. The standard deviation of the percentage change in the real exchange rate is employed to measure the exchange rate volatility. Cointegration and error-correction models are used to obtain the estimates of the cointegrating relations and the short-run dynamics, respectively. The results obtained in this paper, on the whole, provide evidence that the real exchange rate volatility has a significant negative effect on real exports.

JEL Classification Codes: F10, F31.

Key Words: Exchange Rate Volatility; Trade Flows.

1. Introduction

The arrival of the flexible exchange rate system in 1973 produced a significant volatility and uncertainty in exchange rates. This started a debate among policy makers and researchers about the impact of exchange rate volatility on trade flows. However, both the theoretical and empirical studies yielded conflicting results about the relationship between exchange rate variability and international trade flows. Although most models of trade argue that exchange rate volatility increases uncertainty and risk and therefore hinders the trade flows, some other studies suggest otherwise.[1] In addition, the issue is mostly examined for developed countries. There are few studies that investigate the relationship for developing countries due

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[1] McKenzie (1999) surveys both the theoretical and empirical studies in the literature. In addition, see, for example, Akhtar and Hilton (1984), Chowdhury (1993), Thursby and Thursby (1987) and Koray and Lastrapes (1989).

mainly to the lack of sufficient time series data. The purpose of this paper is to close this gap and empirically examine the effect of exchange rate volatility on Turkey's real exports to the United States and to its three major trading partners in the European Union: Germany, France, and Italy.

In estimating these effects, the approach followed here is very close to that of in Kumar and Dhawan's (1991)'s paper, which investigates the topic by estimating an export function for each of the Pakistan's major trading partners in the developed world. This paper is different from the previous papers in several ways. First, this study pays special attention to the specification of the volatility estimator. Two separate versions of exchange rate volatility, the standard deviation of the percentage change in the real exchange rate and the variance of the real exchange rate around its predicted trend, are computed. Having confirmed that both of the versions sufficiently measure the volatility of real exchange rates, the version of the standard deviation of the percentage change in the real exchange rate is employed as one of the independent variables that estimate real exports. Second, this study extends the study by Kumar and Dhawan (1991) in two ways. First, their study fails to recognize that real exports and some of its determinants, i.e., importing country's real income or relative prices, are potentially non-stationary variables. This study focuses upon the nature of non-stationarities apparent in various time series across countries. Second, even though Kumar and Dhawan's paper examines the possibility of lagged relationship between the volume of exports and its various determinants, it does not consider the short run dynamics. Using both cointegration and error-correction techniques, this paper explicitly takes into account the possibility of such a lagged relationship and investigates the long-run relationship between exports and its determinants and also considers the short run dynamics by which exports converge on their long-run equilibrium values. Thus, this work mainly presents additional evidence on the effect of exchange rate volatility on trade flows in the context of developing countries and extends the work of Kumar and Dhawan.

The remainder of the paper is organized as follows. Section 2 presents the specification of the empirical model. Section 3 describes data sources and variable definitions. Section 4 discusses the empirical results. Conclusions are drawn in the last section.

2. Model Specification

Previous studies that have investigated the impact of exchange rate volatility on trade have estimated separate trade equations with exchange rate volatility as one of the explanatory variables in each equation. Using a traditional export demand function with an addition of a measure of exchange rate volatility, the long-run equilibrium export demand can be written as (McKenzie 1998, Chowdhury 1993):

$$\ln X_t = \lambda_0 + \lambda_1 \ln Y_t^f + \lambda_2 \ln P_t + \lambda_3 \ln V_t + u_t \quad (1)$$

where X_t is real exports at time t, Y_t^f is a measure of real foreign economic activity at time t, P_t represents the bilateral real exchange rate at time t, and V_t is the measure of exchange rate volatility at time t. One would expect that increases in real GDP of trading partners result in a greater volume of exports to those partners. In addition, the real exchange rate depreciation (an increase in the level of directly quoted exchange rate) may lead to an increase in exports due to the relative price effect. As explained in the introduction, the relationship between the volatility of the real exchange rate and the real exports is ambiguous. Thus, it is expected that λ_1 and $\lambda_2 > 0$ and $\lambda_3 <$ or >0.

Since this study focuses on the coefficient λ_3, the expectations for the sign of λ_3 are explained in somewhat more detail. The impact of exchange rate volatility on trade volume is ambiguous from a theoretical point of view. The standard theoretical argument that exchange rate volatility may hinder the flow of international trade centered on the notion that exchange rate volatility represents uncertainty and will impose costs on risk averse commodity traders. A number of authors, such as Ethier (1973), Hooper and Kohlhagen (1978), and Wolf (1995) illustrate, in the context of theoretical models, that exchange rate volatility might hamper trade. Contrary to this view, Franke (1991), De Grauwe (1988), and Giovannini (1988) have developed models, which show that exchange rate volatility may actually stimulate trade flows. For example, De Grauwe (1988) argues that if producers are sufficiently risk averse, an increase in exchange rate risk raises the expected marginal utility of export revenue and induces them to export more.

Numerous empirical studies reflect this ambiguity in the theoretical literature. Aristotelous (2001) and McKenzie (1998) find no firm evidence for the relationship between exchange rate volatility and trade. Akhtar and

Hilton (1984), Kenen and Rodrik (1986), Koray and Lastrapes (1989), and Chowdhury (1993), inter alia, provide evidence in support of the view that the volatility of exchange rates reduces the volume of international trade. On the other hand, McKenzie and Brooks (1997), and Klein (1990) find some evidence for a positive effect of exchange rate volatility on trade flows.

The same conflicting evidence for the relationship between exchange rate volatility and trade exists with regard to developing countries. While some studies such as by Kumar and Dhawan (1991), Arize et al., (2000) and Doroodian (1999) found a negative relationship, the study by Warner and Kreinin (1983) failed to report any firm relationship between export flows and exchange rate volatility.

Different statistical measures of exchange rate variability have been used in the literature: Specifically, they are:1- squared residual from ARIMA process, 2-difference between previous forward and current spot rates, 3-Gini mean difference coefficient, 4- dummy for exchange rate regime, 5- the volatility measure generated from generalized autoregressive conditional heteroscedastic (GARCH) process.[2] In this paper, since there is no single theoretically correct measure of variability, following Kenen and Rodrik (1986) and Thursby and Thursby (1993), two different versions of the exchange rate volatility are computed. In the first measure, the variance of the real exchange rate around its predicted trend is estimated from the equation:

$$\ln E^{i}_{jt} = \psi^{i}_{0} + \psi^{i}_{1}\, t + \psi^{i}_{2}\, t^{2} + u^{i}_{jt} \tag{2}$$

where E is the real exchange rate. The trend equations for each month are estimated for the twelve preceding months and the residual variance is assumed to be the volatility measure for that month. In the second measure, the standard deviation of the percentage change in the real exchange rate for the twelve preceding months is computed as the volatility measure for that month. Even though the assumptions of the volatility measures are different from each other, the comparison of the two versions for each country reveals very strong correlations indicating that both of the versions adequately measure the variability of real exchange rates.[3] Thus, only the second

[2] Even though GARCH models are becoming increasingly popular to estimate volatility, it does not work very well in finite samples. See McKenzie (1998).

[3] The least correlation coefficient between two measures is 0.89 indicating almost perfect positive association for each country.

version of the volatility measure is used as one of the explanatory variables in the real export equations.

3. Data and Variable Definitions

Turkey's exports to the United States and its three major trading partners in the European Union- Germany, France and Italy- are examined for the period between 1990:01 and 2000:12. Even though Turkey adopted more liberal policies for the domestic economy after 1981, it did not fully liberalize its exchange rate policy until 1988. Before 1988, Turkey implemented adjustable peg policy in which the Turkish lira (TL) was daily adjusted in the form of devaluations. Thus, the sample period from 1990:01 to 2000:12 is chosen to minimize the specification problems stemming from the change in exchange rate policies of Turkey.

Data were obtained from the IMF's *International Financial Statistics*, OECD *Main Economic Indicators* and OECD *Monthly Statistics of International Trade*. Nominal exports of Turkey to each trading partner were defined in the United States Dollars and are deflated by the U.S.A. consumer price index to define them in real terms.

Economic theory suggests that income in an importing country is a major determinant of a nation's exports. Since monthly GDP statistics are rarely found, a proxy for GDP is required. The OECD publishes a monthly measure called the index of industrial production, which is used as a proxy for the trend of GDP.[4]

Second explanatory variable in the export equation measures competitiveness ($P_{t)}$, where P_t is the bilateral real exchange rate between Turkey and its trading partners. Like Kenen and Rodrik (1986), we work with volatility in the real exchange rate. Real exchange rates (REAL) were derived from monthly nominal exchange rates for the Turkish Lira against each country's currency (NOM), a monthly Turkish consumer price index (CPITR), and a monthly foreign country consumer price index (CPIFR).[5]

$$\text{REAL} = (\text{NOM}*\text{CPIFR})/\text{CPITR} \tag{3}$$

[4] McKenzie and Brooks (1997) and Doroodian (1999) follow the same approach.
[5] The literature has commonly used the CPI to deflate the nominal exchange rates (see, for example, Thursby and Thursby; 1987 and Caporale and Doroodian; 1994)

CPI indices are from OECD *Main Economic Indicators* database. The monthly bilateral nominal exchange rates data for the United States for the whole period are from OECD *Main Economic Indicators*. For Germany, France and Italy, the monthly nominal exchange rates were pieced from two different sources. The data for the periods between 1990:01 and 1998:12 are from IMF's *International Financial Statistics*. For the periods between 1999:01 and 2000:12, daily data were obtained from *FXTOP* company's currency converter site and then converted to a monthly basis by taking average of the daily data.[6]

4. Empirical Results

This section reports the estimates for Turkey's export functions. In order to detect the long-term co-movement among the variables included in the equation (1), the cointegration procedure developed in Johansen (1991) and Johansen and Juselius (1990) is employed. For each of Turkey's trading partners in the sample, an error-correction model for real exports is developed.

4.1 Cointegration Analysis

Before estimating the cointegration parameters, the order of integration of each series should be examined. The order of integration of the individual time series is determined using the augmented Dickey-Fuller (ADF) test recommended by Engle and Granger (1987). The ADF test statistics for a variable y_t is given by the t statistics on the estimated coefficient μ, in the following regression:

$$\Delta y_t = \alpha + \beta_t \, \text{trend} + \mu y_{t-1} + \sum_{j=1}^{k} \delta_j \Delta y_{t-j} \qquad (4)$$

where y is relevant variable and k is determined by using the minimum value of the Akaike Information Criterion. The specification is then used to test:

$$H_0 : \mu = 0 \text{ and } H_1 := \mu < 0.$$

[6] *FXTOP* company's currency converter site, http://fxtop.com/en/historates.php3 (05-01-2002), complies with the European Union Council Regulation, 97/1103.

The regression (4) is estimated with and without trend variable. Irrespective of the country considered, the empirical results in Table 1 and Table 2 indicate that the variables included in this study are integrated of order one.

In applying the Johansen procedure, one needs to specify the number of lags in each cointegrating equation and to choose the one that seems most plausible for the data in hand from the five possible specifications in Johansen (1995). Akaike Information Criterion (AIC) is used to determine the number of lags in the cointegrating equations. The AIC results indicate that the optimum lag length is one for all countries. An examination of series for each country suggested us to use the specification that the cointegration test assumes deterministic trend in data. In addition, the likelihood ratio test allows the specification for all countries that the cointegrating equations have only intercepts.

Table 3 reports the results from the Johansen likelihood ratio tests for cointegration. The two common likelihood-ratio tests, the trace and the maximum eigenvalue (λ-max) tests, are used to determine the number of cointegrating relations in non-stationary time series. For λ-max and trace statistics, the null hypothesis is that there are r or fewer cointegration vectors, whereas the alternative hypotheses are $r+1$ and at least $r+1$ cointegrating vectors for the λ-max and trace statistics, respectively.

Starting with the λ-max results, the null hypothesis of $r = 0$ (no cointegration) is rejected in favor of the alternative hypothesis $r = 1$ in each country. On the other hand, the null hypotheses of $r \leq 1$, $r \leq 2$ and $r \leq 3$ cannot be rejected in favor of the alternative hypotheses of $r = 2$, $r = 3$ and $r = 4$, respectively. These results indicate the presence of only one cointegrating relationship for each country.

Similar conclusions are obtained from the trace test results. The null hypothesis of $r = 0$ is rejected in favor of $r \geq 1$ in each country. Furthermore, the null hypotheses of $r \leq 1$, $r \leq 2$ and $r \leq 3$ cannot be rejected in favor of the alternative hypotheses of $r \geq 2$, $r \geq 3$ and $r \geq 4$, respectively, for all countries. The trace test results indicate the presence of only one cointegrating relationship for each country. The results from both of the two tests suggest that there is a long-run equilibrium relationship among real exports, foreign country's income, real exchange rate and exchange rate volatility for all countries in our sample.

The cointegrating vectors, which are normalized with respect to the real exports, together with their respective t-values, are given in Table 4. The results of this normalization yield estimates of long-run elasticities. Foreign economic activity ($Y_t^{f)}$ is positively related to real exports (X_t) for all sample countries and the coefficients are statistically significant at the 1 % level for three countries and at the 10 % level for one country. It must be noted that the values for income elasticities are consistent with estimates found in other studies such as Arize et al. (2000). The estimated competitiveness variable, the real exchange rate, (P_t), has an expected positive sign and it is significant for all countries at the 1% level. An appealing aspect of the results is that the volatility measure is negative for all countries and it is significant at the 1% level for France and Germany and at the 10% level for the United States. These results provide strong evidence that exchange rate volatility has a negative and significant long-run effect on real exports.

4.2 Error-Correction Model

Based on the theorem developed in Engle and Granger (1987), the existence of a cointegrated relationship among a set of *I* (1) series implies the following dynamic error correction representation of the data:

$$\Delta \ln X_t = \alpha_0 + \alpha_1 ECT_{t-1} + \phi_i \sum_{i=1}^{k} \Delta \ln X_{t-i} + \beta_i \sum_{i=0}^{k} \Delta \ln Y^{f}_{t-i} + \lambda_i \sum_{i=0}^{k} \Delta \ln P_{t-i} +$$

$$\gamma_i \sum_{i=0}^{k} \Delta \ln V_{t-i} + \varepsilon_t \qquad (5)$$

where ECT_{t-1} is the lagged error correction term and is the residual from the cointegrating regression equation (1). It should be noted that the error correction term, ECT ~ *I* (0), captures the adjustment toward the long-run equilibrium. The coefficient α_1 represents the proportion of the disequilibrium in real exports in one period corrected in the next period. The equation (5) is estimated with a general specified lag structure for all the variables in the equation (1), a constant term and one-lagged error-correction term. The lag length for the VAR models for each country is determined using the likelihood ratio test.

The estimation results of the VAR are summarized in Table 5. Each estimated models fulfills the conditions of serial noncorrelation, homoscedasticity and no specification errors. The adjusted R^2's are ranging from 0.28 to 0.36. The low R^2 values are due to the reason that regressions are based on the first differences in variables.

For all countries, the error correction term's coefficient is negative and statistically significant at the 1% level further confirming that the variables are cointegrated. The magnitudes of the error correction terms indicate the change in real exports per month that is attributed to the disequilibrium between the actual and equilibrium levels. The adjustment of speed to the last period's disequilibrium varies across countries substantially implying that, for example, while 61% of the adjustment occurs in one month for the Italy, 44% of the adjustment occurs in one month for Germany.

The primary interest in this section is that the estimated volatility measure is negative and significant at the 10% level for Germany and negative and insignificant in for the Unites States and France and positive and insignificant for Italy. Although the exchange rate volatility has a significant negative long-run effect on real export demand for three countries in the sample, exchange rate volatility has a significant negative short run effect on real export demand only for Germany. The exchange rate risk seems less of a factor in explaining export demand for the other countries in the short run.

Finally, although the expected sign of the foreign country's income and the real exchange rate has occurred in the long run for all countries, the coefficient for foreign country's income is positive and significant only for Italy and insignificant for the rest of the countries and the real exchange rate is negative and significant for Italy and insignificant for the other countries.

Each of the real exports is evaluated by examining their structural stability using the Chow test over the 1990:01-2000:12. The potential breakpoint is the 1994:4 in which Turkey experienced major economic crisis which may lead to structural changes. The test statistics for each country is the F-statistic, which is based on the comparison of the restricted and unrestricted sum of squared residuals. The results indicate in Table 5 that we cannot reject the null hypothesis of structural stability and thus confirm the stability of each estimated error correction model for each country.

5. Summary and Conclusions

This paper reexamines the impact of exchange rate volatility on the demand for real exports in the context of a multi-variate error-correction model. The model is estimated for Turkey's real exports to each of its three major trading partners in the European Union and the United States for the period between 1990:01 and 2000:12. We estimate two exchange rate volatility measures; the variance of the real exchange rate around its predicted trend and the standard deviation of the percentage change in the real exchange rate. Having confirmed that both of the versions adequately measure the variability of real exchange rates, the version of the standard deviation of the percentage change in the real exchange rate is used as a proxy for exchange rate risk. The empirical results based on cointegration analysis show that real exports are cointegrated with foreign income, real exchange rate and exchange rate volatility.

Our results concerning the effects of exchange rate volatility on real exports suggest that the long-run relationship between Turkey's real exports and its exchange rate volatility is negative and statistically significant for Germany, France and the United States. In addition, the exchange rate volatility has negative short-run effects on real exports to Germany. For the rest of the countries, the short-run impact of the exchange rate volatility is statistically insignificant. Utilization of forward exchange markets to fully hedge exchange rate risk may have made exchange rate volatility less of a factor in explaining real exports to these countries in the short-run.

These results, on the whole, provide uniform evidence on the effect of the exchange rate volatility on real exports. This finding supports those who point out that exchange rate volatility have a negative impact on trade. In addition, the Chow test results show the absence of parameter instability in the estimated models.

Finally, to the extent that one can generalize from Turkey's experience, policy makers in developing countries should consider both the existence and the degree of exchange rate volatility and notice the likely impact of the exchange rate volatility for each trading partner in implementation of trade policies.

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Table 1: Augmented Dickey Fuller Unit Root Test Results (without trend)

Countries	Germany	Italy	United States	France
$\ln X_t$	-2.38	-1.68	-0.38	-0.27
$\ln Y_t^f$	0.29	0.43	0.44	1.39
$\ln P_t$	-1.06	-1.92	-2.79	-1.18
$\ln V_t$	-1.75	-1.51	-1.99	-2.01
$\Delta \ln X_t$	-8.44*	-8.73*	-7.21*	-8.05*
$\Delta \ln Y_t^f$	-9.43*	-11.40*	-4.35*	-8.29*
$\Delta \ln P_t$	-4.36*	-7.43*	-7.33*	-4.31*
$\Delta \ln V_t$	-6.52*	-6.91*	-6.49*	-6.34*

Notes: * shows rejection of the null hypothesis of a unit root at the 1% level. The MacKinnon critical value for the 5% level is-2.88 and 1% level is -3.48. The lag order for the series was determined by the Akaike Information Criterion. The symbol Δ is the first difference.

Table 2: Augmented Dickey Fuller Unit Root Test Results (with trend)

Countries	Germany	Italy	United States	France
$\ln X_t$	-2.20	-2.62	-2.99	-2.08
$\ln Y_t^f$	-0.61	-2.38	-3.22	-1.46
$\ln P_t$	-1.19	-3.10	-3.06	-1.51
$\ln V_t$	-2.92	-2.65	-3.26	-3.14
$\Delta\ln X_t$	-7.18*	-8.71*	-8.70*	-8.48*
$\Delta\ln Y_t^f$	-9.51*	-11.63*	-4.45*	-8.81*
$\Delta\ln P_t$	-4.50*	-7.43*	-7.30*	-4.41*
$\Delta\ln V_t$	-6.54*	-6.91*	-6.46*	-6.36*

Notes: * shows rejection of the null hypothesis of a unit root at the 1% level. The MacKinnon critical value for the 5% level is-3.44 and 1% level is -4.03. The lag order for the series was determined by the Akaike Information Criterion. The symbol Δ is the first difference.

Table 3: Johansen Cointegration Test Results

		λ-max Statistics				Trace Statistics			
	H_0	r = 0	r ≤ 1	r ≤ 2	r ≤ 3	r = 0	r ≤ 1	r ≤ 2	r ≤ 3
Country	H_1	r = 1	r = 2	r = 3	r = 4	r ≥ 1	r ≥ 2	r ≥ 3	r ≥ 4
The US		44.06	16.73	8.38	0.02	69.73	25.13	8.40	0.02
Italy		43.28	7.12	4.36	0.07	54.83	11.55	4.43	0.07
Germany		33.12	12.15	7.06	0.44	52.72	19.65	7.50	0.44
France		43.18	12.18	6.38	1.04	62.78	19.60	7.42	1.04
CV (5%)		27.07	20.97	14.07	3.76	47.21	29.68	15.41	3.76

Note: *r* denotes the number of cointegrating vectors. The critical values are for the 5% level of significance reported by Osterwald-Lenum (1992).

Table 4: Estimates of the Cointegrating Relationships

Country	Normalized Cointegrating Vector
The United States	$\ln X_t = -27.04 + 1.73 \ln Y_t^f + 2.24 \ln P_t - 0.15 \ln V_t$ (3.60) (4.39) (-1.66)
France	$\ln X_t = -8.30 + 2.17 \ln Y_t^f + 0.31 \ln P_t - 0.13 \ln V_t$ (6.20) (2.58) (-6.50)
Italy	$\ln X_t = -14.91 + 3.79 \ln Y_t^f + 0.65 \ln P_t - 0.02 \ln V_t$ (6.01) (2.82) (-0.40)
Germany	$\ln X_t = -5.45 + 0.89 \ln Y_t^f + 0.72 \ln P_t - 0.14 \ln V_t$ (1.64) (4.00) (-4.66)

Note: The numbers in parentheses beneath the estimated coefficients are t-statistics.

Table 5: Turkey's Real Export Equations

Country	United States	Italy	Germany	France
ECT_{t-1}	-0.54 (-5.82)	-0.61 (-6.66)	-0.44 (-5.42)	-0.68 (-6.53)
ΔlnX_{t-1}	-0.17 (-2.09)	-0.004 (-0.05)	-0.16 (-1.93)	-0.07 (-0.86)
ΔlnY^{f}_{t}	-3.35(-0.87)	2.15(1.98)	0.66 (0.76)	0.56 (0.39)
ΔlnP_t	0.10 (0.24)	-0.73 (-2.22)	-0.26 (-1.12)	-0.36 (-1.45)
ΔlnV_t	-0.06 (-0.60)	0.01 (0.11)	-0.13 (-1.86)	-0.02(-0.40)

Summary Statistics

	United States	Italy	Germany	France
$\bar{R}^2$	0.34	0.33	0.28	0.36
SC	χ^2 (1) =3.03 (0.08)	χ^2(1) =1.95 (0.16)	χ^2 (1) =0.14 (0.69)	χ^2(1)=1.35 (0.24)
ARCH	χ^2 (2) =0.28 (0.86)	χ^2(2) =1.61 (0.44)	χ^2 (2) =0.29 (0.86)	χ^2 (2) =0.12 (0.94)
RESET	F(2, 123)=0.63 (0.53)	F(2, 123)=2.68 (0.07)	F(2, 123)=0.94 (0.39)	F(2, 123)=0.40 (0.66)
CHOW	F (6,119)=1.33 (0.24)	F (6,119)= 1.07 (0.38)	F (6,119)= 0.72 (0.63)	F (6,119)=1.13 (0.34)

Notes: The numbers in parentheses after the coefficients are the *t* statistics, as those of after the summary statistics are the *p* values. The symbol Δ is the first difference. ARCH [χ^2(q)] is the chi-square test for autoregressive conditional heteroscedasticity.

SC [χ^2 (q)] is the qth order Breusch-Godfrey LM test statistic for serial correlation.
RESET[F(q, T-k)] is the qth order Ramsey's RESET test statistics.
CHOW is the Chow breakpoint test for structural stability.