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Exchange rate uncertainty in Turkey and its impact on export volume

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Abstract

This paper investigates the impact of real exchange rate volatility on Turkey's exports to its most important trading partners using quarterly data for the period 1982 to 2001. Cointegration and error correction modeling approaches are applied, and estimates of the cointegrating relations are obtained using Johansen's multivariate procedure. Estimates of the short-run dynamics are obtained through the error correction technique. Our results indicate that exchange rate volatility has a significant positive effect on export volume in the long run. This result may indicate that firms operating in a small economy, like Turkey, have little option for dealing with increased exchange rate risk.

1. Introduction

Since the advent of generalized floating by industrial countries in 1973, an increasing number of developing countries, including Turkey, abandoned fixed exchange rate regimes and adopted various forms of flexible exchange rate arrangements. There is concern over the flexible exchange rate system, however, because of the high degree of volatility and uncertainty of exchange rate movements and the effect of such volatility on international trade. The conventional argument is that increased uncertainty from high volatility in the exchange rate can affect international trade, and may therefore reduce the advantages of worldwide specialization. Nonetheless, there is no real consensus about the effects of exchange risk on trade volume, either theoretically and empirically.

On the theoretical front, Baron (1976) and Giovannini (1988) produced models that show how an increase in exchange rate volatility may not necessarily affect the level of trade. Franke (1991), De Grauwe (1988), Sercu and Vanhulle (1992) and Dellas and Zilberfarb (1993) have produced models that provide a counter-intuitive result: they show how increased exchange rate volatility may actually lead to greater levels of trade. For example, De Grauwe (1988) has argued that high risk could actually lead to increased exports. Exchange rate volatility unambiguously reduces the total utility to be derived from exporting, but would result in increased exports if the marginal utility of exporting increased (the firm is assumed to be engaged in both the domestic market and the export market, and allocating output optimally between both markets). Crucial to this result is the idea that the degree of risk aversion is not constant. If it were constant, then exchange rate volatility would unambiguously reduce the level of exports, as exporting becomes a relatively less attractive activity (substitution effect). There would be no income effects. Alternatively, if the degree of risk aversion increases with shrinking income, then the income effect will lead exporters to export even more in response to increased exchange rate volatility, in order to avoid the utility depression effect of a large reduction in their export earnings. More recently, Gagnon (1993), Broll (1994), and Wolf (1995) have produced models that support the idea that exchange rate volatility acts to the detriment of international trade. Sercu and Uppal (1997) have presented models that show how volatility may impact either positively or negatively on trade depending on the underlying assumptions. Hence, the influence of exchange rate volatility on trade volume is ambiguous from a theoretical point of view.

Reflecting the theoretical debate, developments in the empirical arena are similarly indecisive. A number of studies test for stationarity of the relevant time series and, in some cases, employ cointegration techniques. Kenen and Rodrik (1986), Peree and Steinherr (1989), Pozo (1992), Chowdhury (1993), Holly (1995), Arize (1995, 1997), Arize *et al*. (2000) and Fountas and Aristotelous (1999), among others, find a statistically significant negative relationship between exchange rate volatility and trade. Asseery and Peel (1991), Doyle (2001) and Bredin *et al*. (2003), however, show evidence of a positive relationship between exchange rate volatility and trade, while Gotur (1985), Koray and Lastrapes (1989) and Bailey *et al*. (1986) were unable to find evidence of any significant effect of exchange rate volatility on trade[1].

[1] McKenzie (1999) provides a comprehensive review of this empirical literature.

Results of the studies on developing economies related to this area also do not unanimously agree with the negative relationship between exchange rate volatility and trade flow. Several studies, for example, Arize *et al*. (2000), and Doroodian (1999), investigate the impact of exchange uncertainty on export volume in developing countries. In these studies, it is assumed that exchange rate risk is more important in developing country trade flows since financial markets for hedging currency risk are not well developed. The main result in these studies is that increases in the volatility of the real effective exchange rate exert a significant negative effect on export demand in both the short-run and the long run. Rose (1990) examines the empirical impact of the real exchange rate on trade balance of a number of developing countries. Using non-structural techniques, the author is unable to find a strong stable effect of the exchange rate on the trade balance.

The effect of exchange rate volatility on international trade and economy is mainly an empirical question. A large number of empirical studies in this literature have been done for the developed countries. This issue is particularly important for countries that switched from a fixed to a flexible exchange rate regime due to the higher degree of variability associated with a flexible exchange rate. While many developing countries have moved to a flexible exchange rate regime within the last two decades, it is surprising that there are only a few studies that analyze the relationship between exchange rate volatility and foreign trade for the developing countries. Hence, the aim of this study is to add to the relatively small stock of evidence on this issue in the context of developing countries by analyzing the impact of exchange rate volatility on Turkey's exports. We examine how exchange rate volatility affects trade by empirically assessing the case of Turkish exports to its nine major trading partners from 1982 to 2001.

On January 24, 1980, the Turkish government announced one of the most important economic stabilization and reform programs in its history. The program included both trade and financial liberalization. The main objective of the financial liberalization program was to increase competition and efficiency of the financial system, while the trade reform program was the core of the liberalization program. The primary objective of trade liberalization was to move towards a liberalized trade regime with flexible exchange rate management. Rapid export growth is a principle objective of these reforms[2].

Adoption of an actively managed, flexible exchange rate system was an important step undertaken in the early 1980s, and it has remained a

[2] See Öniş and Riedel (1993).

central instrument in the trade liberalization program[3]. The international trade performance of a small open economy, such as Turkey, plays a central role in the economic health of the country. The share of Turkish merchandise exports in GDP has increased dramatically in recent years (from 15.63% in 1987 to 34.5% in 2001)[4], thus rendering the economy more open than before and more dependent on foreign markets. Hence, policies designed to enhance export performance are of increasing importance to national economic welfare. Correct policy decisions depend on having relevant information on the factors that affect the level of exports. We aim to provide information on one of these factors, the volatility of exchange rates.

The rest of the paper is organized as follows. In section 2, we examine the specification of our empirical model and data. Section 3 reports the empirical results. The paper's concluding remarks are provided in section 4.

2. Empirical model and data

Drawing upon the empirical literature in this area, the standard long run relationship between real exports, the level of real activity, competitiveness, and exchange rate volatility is specified as follows (see for example, Asseery and Peel, 1991; Chowdhury, 1993; Arize, 1995, 1997):

$$\ln X_t^* = \beta_0 + \beta_1 \ln Y_t + \beta_2 \ln P_t + \beta_3 \ln V_t + \varepsilon_t \qquad (1)$$

where X_t^* denotes the desired volume of a country's export goods; Y_t is a measure of real foreign income; P_t stands for relative prices; and V_t represents exchange rate volatility.

Economic theory suggests that income in trading partner countries is a major determinant of a nation's export performance. If foreign income

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[3] Turkey has adopted several different exchange rate systems since the stabilization program of 1980. The prices of the foreign currencies were determined under the flexible exchange rate system in the form of a crawling band for the period 1981-1988, and in the form of a managed float between 1989 and 1993. The managed float was transformed into a dirty float in 1994. After a 5-year period of a managed float regime between 1994 and 1999, a crawling peg system was adopted under a standby agreement with the IMF in 1999 in order to control the chronic and high inflation rates. The system had a short life due to the financial crisis of 2001. Foreign exchange rates are determined under a flexible exchange rate system since February 2001.

[4] The Turkish economy experienced two major financial crises in recent years. These financial crises affected the real sector deeply, leading to negative growth rates in GDP. The Turkish Lira depreciated by 125% in 1994 and 100% in 2001 against the US dollar. The share of exports in GDP increased due to the negative growth rate in GDP.

rises, the demand for exports will rise; so β_1 is expected to be positive. If relative prices rise, the demand for exports will fall, and β_2 is expected to be negative. Most empirical work treats exchange rate uncertainty as a risk: higher risk leads to higher cost for risk-averse traders and also less trade. As Bailey, Tavlas, and Ulan (1986) point out, traders may anticipate future exchange rate movements better than the average exchange market participant, and gains from this knowledge could offset the risk of exchange rate uncertainty. Moreover, if the exchange rate volatility is due to fundamentals, efforts by the authorities to reduce it by means of exchange controls or other restrictions on trade could be more harmful to trade and could reduce it more. Hence, the effect of exchange rate uncertainty of export demand cannot be determined *a priori***,** but is an empirical matter.

To make equation (1) estimable, we need to replace the desired export demand with actual (observable) levels (i.e., $X_t^* = X_t$). A series defined as the real foreign income is constructed by taking the weighted average of the GDP series of Turkey's nine most important trading partners, namely Germany, the United States, United Kingdom, Italy, France, the Netherlands, Spain, Belgium, and Greece[5]. Weights are calculated as the sum of exports from Turkey to each country as a share of Turkey's total exports to these countries[6]. The individual GDP series is converted to a common currency—US dollar—for aggregation purposes. The second explanatory variable in the export equation measures competitiveness where P_t is defined as the ratio of export price index of Turkey to the weighted average of export price indices of major trading partners. The weights are identical to those used in the construction of the income variable. All quarterly data are taken from the International Financial Statistics tape of the IMF and the Central Bank of the Republic of Turkey electronic data delivery system. The sample period runs from the first quarter of 1982 to the fourth quarter of 2001.

It is necessary to derive a measure of exchange-rate uncertainty. In this paper, we use the moving standard deviation of the growth rate of the exchange rates[7]. This proxy is constructed as follows:

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[5] Choosing the weighted average of the income levels of the most important trading partners is standard procedure in the literature (see Chowdhury, 1993).

[6] About 60% of Turkey's exports go to these countries.

[7] Recently, several authors model exchange rate volatility using an ARCH approach. McClain *et al*. (1996) suggest that 300 observations is a threshold value for estimating a reliable ARCH model. Since we have only 80 observations we have not taken this route.

$$V_t = \left[(1/m) \sum_{i=1}^{m} (\ln R_{t+i-1} - \ln R_{t+i-2})^2 \right]^{0.5} \quad (2)$$

where R is the real effective exchange rate and m, the order of the moving average, is set to equal to 5[8]. The real effective exchange rate is calculated by the weighted average of the exchange rate-adjusted relative prices (unit export values) where the trade weights are the ones used in creating foreign income and relative prices. Details of the calculation of the real effective exchange rate is provided in Appendix 1. This measure of exchange rate volatility is adopted by Kenen and Rodrik (1986), Koray and Lastrapes (1989), and Chowdhury (1993).

Although almost all studies in this literature use real effective exchange rates in calculating volatility, there has been some conflicting arguments as to whether exchange rate uncertainty is better measured by nominal or real exchange rate volatility[9]. To avoid such a confliction, three different measures of volatility are used to identify how the results vary across volatility measures[10]. Following equation (2) the measures of volatility are as follows:

a) moving sample standard deviation of the real effective exchange rate;
b) moving sample standard deviation of the nominal effective exchange rate (using the US dollar);
c) moving sample standard deviation of the nominal effective exchange rate (using the German DM).

For simplicity, throughout the analyses these measures are referred to V, VUS and VDM respectively[11].

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[8] Our main results are robust to alternative choices of the lag length.

[9] Akhtar and Hilton (1984) have reported significant trade flow effects of nominal exchange rate volatility, while Kenen and Rodrik (1986), Arize (1995,1997) and Arize *et al*. (2000), among others, have found significant trade flow effects of real exchange rate volatility. Thursby and Thursby (1987), Lastrapes and Koray (1990) and Doyle (2001), however, find a significant impact for volatility on exports when using both nominal and real exchange rates. We should mention that most of the empirical studies in this literature use real exchange rates in the calculation of volatility. In theoretical studies (see, for example, De Grauwe, 1994), the real exchange rate is used as a parameter in the profit function. Since real exchange rates include relative prices (CPI or WPI), it provides information on whether the domestic currency is overvalued or undervalued, which also affects volatility.

[10] We wish to thank an anonymous referee who pointed out this issue

[11] Germany and US are two of the major trading partners. Although evidence for other countries indicates that trade is predominantly invoiced in the exporter's currency (see Hooper and Kohlhagen, 1978), data from the Undersecretariat of Foreign Trade indicate that, on average, 85% of all export contracts were written against the US dollar and DM between 1989 and 1998. After the EURO was introduced in 1999, about 92% of all export contracts were written against the US dollar and EURO.

2.1. The specific model

Many studies have employed the simple stock adjustment mechanism whereby the entire adjustment is represented by adding a lagged dependent variable as a regressor to allow for the adjustment of export demand to changes in the regressors. However, several researchers have criticized this stock adjustment structure because of its restrictive assumptions. Moreover, such an equation is subject to estimation problems such as the 'spurious regression phenomenon'. This phenomenon refers to the possibility that inferences based on ordinary least-squares parameter estimates in such regressions are invalid because the usual *t*- and *F*- ratio test statistics do not converge to their limiting distribution as the sample size increases (Arize, 1995). Their use in that case generates spurious inferences if the levels of the nonstationary variables included in equation (1) are not cointegrated.

Recent advance in cointegration and dynamic modeling techniques suggest some statistical procedures for addressing these issues. We assume that equation (1) is the cointegrating equation to establish the long run equilibrium relationship among the variables. To identify the short run dynamic specification, we employ the following error correction model (ECM):

$$\Delta \ln X_t = \beta_0 + \sum_{i=1}^{n} \beta_{1i} \Delta \ln X_{t-i} + \sum_{i=0}^{n} \beta_{2i} \Delta \ln Y_{t-i} + \sum_{i=0}^{n} \beta_{3i} \Delta \ln P_{t-i} + \sum_{i=0}^{n} \beta_{4i} \Delta V_{t-i} + \beta_5 EC_{t-1} + \varepsilon_t \qquad (3)$$

where EC_{t-1} is the lagged error correction term and is the residual from the cointegrating regression equation (1). If the variables employed in equation (1) are cointegrated, then the error correction form in equation (3) exists.

The modeling strategy adopted in this study involves three steps:

a. determining the order of integration of the variables by employing Augmented Dickey-Fuller (ADF) unit-root tests;
b. if the variables are integrated of the same order, we test for cointegration by applying the Johansen-Juselius (1990) approach[12]; and,
c. if the variables are cointegrated, we can specify an error correction model and estimate it using standard methods and diagnostic tests.

[12] It is now well established that the Johansen (1988) cointegration procedure is superior to the residual-based model (see Gonzalo, 1994).

3. Empirical results

As a preliminary step to cointegration analysis, the stationarity of each of the variables was tested using Augmented Dickey-Fuller (ADF) and Phillips-Perron tests. The ADF test consists of regressing each series on its lagged value and lagged difference terms. The ADF tests results are shown in Table 1. The results suggest that all variables in equation (1) are nonstationary in their levels and they are integrated of order one[13]. Therefore, we can proceed to the cointegration tests.

3.1. Cointegration test

The Johansen test statistics (trace and maximum eigenvalue) are used to identify the presence of common stochastic trends at each measure of volatility. To determine the number of cointegrating vectors, a vector autoregression (VAR) was used, with lags of each variable chosen on the basis of the Akaike and Schwarz information criteria[14]. Table 2 reports both maximum eigenvalue and trace statistics.

Starting with the null hypothesis of no cointegration ($r = 0$) among the variables, the maximal eigenvalue statistic is 41.5, which is above the 95% critical value of 28.1. Hence it rejects the null hypothesis in favor of the general alternative ($r = 1$). As is evident in Table 2, 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 among the four variables.

For the trace test results, we obtain similar conclusions when the null hypothesis of $r = 0$ is tested against the alternative hypothesis of $r \geq 1$. But the test fails to reject the null hypotheses of $r \leq 1$, $r \leq 2$, and $r \leq 3$. In the case of nominal volatility, we obtained very similar results, indicating the presence of only one cointegrating relationship among the

[13] The results of the Phillips-Perron unit root tests are similar and are not reported here. One criticism of unit root testing is that a stationary series subject to a structural break can look like a nonstationary series. If the structural break (or breaks) is not taken into account, the unit root test leads to false nonrejection of the null hypothesis of nonstationarity. Therefore, too often series are falsely found to be nonstationary. Since the Turkish economy witnessed one of the most important financial crises in its history in 1994, there might be a structural break in that year. To check the effect of a possible structural break on unit root tests due to the financial crisis we followed an approach suggested by Perron (1989). In this approach, a single breakpoint is assumed, which is incorporated into the regression model. We used three tests (with trend and without trend) suggested by Perron (1989) to determine the order of integration of the variables. Our results suggest the presence of unit roots in all variables used in the analysis.

[14] Four lags were chosen for each specification (e.g., LV, LVUS and LVM).

Table 1
ADF Unit Root Tests for Stationarity

Variables	Level/first difference	Without trend	With trend	Conclusion
LX	Level	-0.929 (4)	-3.337 (4)	I(1)
		(-3.520)	(-3.161)	
	First difference	-4.564 (4)*	-4.639 (4)*	I(0)
		(-3.521)	(-4.087)	
LP	Level	-2.383 (3)	-2.030 (3)	I(1)
		(-3.519)	(-4.083)	
	First difference	-4.799 (3)*	-4.996 (3)*	I(0)
		(-3.520)	(-4.085)	
LGDP	Level	-1.958 (3)	-2.036 (3)	I(1)
		(-3.518)	(-4.084)	
	First difference	-3.872 (3)*	-5.387 (3)*	I(0)
		(-3.520)	(-4.085)	
LV	Level	-2.796 (4)	-2.557 (4)	I(1)
		(-3.520)	(-4.085)	
	First difference	-4.824 (4)*	-5.045 (4)*	I(0)
		(-3.521)	(-4.087)	
LVUS	Level	-0.766 (4)	-2.987 (4)	I(1)
		(-3.525)	(-4.093)	
	First difference	-4.891 (4)*	-5.090 (4)*	I(0)
		(-3.527)	(-4.095))	
LVDM	Level	-0.276 (2)	-2.827 (2)	I(1)
		(-3.522)	(-4.089)	
	First difference	-5.458(2)*	5.520 (2)*	I(0)
		(-3.524)	(-4.091)	

*Indicates significance at 1% level. Lag lengths are chosen based on the likelihood ratio, the Akaike information criteria (AIC), and Schwarz information criteria (SC).
Note: Figures in parentheses underneath the estimated coefficient are the critical values at 1% (MacKinnon, 1991).

four variables. Hence, overall our findings suggest that there is a long run equilibrium relationship among real exports, foreign income, relative price, and exchange rate volatility. Our results agree with the findings of Arize, *et al*. (2000) in their empirical study on thirteen developing countries. With the same econometric framework, they observe a long run relationship among the variables for the countries in their sample for the period 1970-1996.

For each measure of volatility, the cointegrating vectors are reported in Table 3. Panel A of the Table indicates cointegrating vectors when we use time-varying moving average of the real effective exchange rate as a

Table 2
Johansen-Juselius Maximum Likelihood Cointegration Tests

Maximum Eigenvalue test				Trace test			
Null	Alternative	Statistic	95% critical value	Null	Alternative	Statistic	95% critical value
A: Results from cointegration tests using LV as a volatility measure							
$r = 0$	$r = 1$	41.48*	28.14	$r = 0$	$r \geq 1$	74.10*	53.12
$r \leq 1$	$r = 2$	18.49	22.00	$r \leq 1$	$r \geq 2$	32.62	34.91
$r \leq 2$	$r = 3$	10.75	15.67	$r \leq 2$	$r \geq 3$	14.13	19.96
$r \leq 3$	$r = 4$	3.38	9.24	$r \leq 3$	$r \geq 4$	3.38	9.24
B: Results from cointegration tests using LVUS as a volatility measure							
$r = 0$	$r = 1$	45.45*	28.14	$r = 0$	$r \geq 1$	77.76*	53.12
$r \leq 1$	$r = 2$	15.78	22.00	$r \leq 1$	$r \geq 2$	32.32	34.91
$r \leq 2$	$r = 3$	11.48	15.67	$r \leq 2$	$r \geq 3$	16.53	19.96
$r \leq 3$	$r = 4$	5.06	9.24	$r \leq 3$	$r \geq 4$	5.06	9.24
C: Results from cointegration tests using LVDM as a volatility measure							
$r = 0$	$r = 1$	33.74*	28.14	$r = 0$	$r \geq 1$	62.79*	53.12
$r \leq 1$	$r = 2$	18.21	22.00	$r \leq 1$	$r \geq 2$	9.06	34.91
$r \leq 2$	$r = 3$	7.39	15.67	$r \leq 2$	$r \geq 3$	0.85	19.96
$r \leq 3$	$r = 4$	3.46	9.24	$r \leq 3$	$r \geq 4$	3.46	9.24

* Indicates significance at 5% level.
Note: r stands for the number of cointegrating vectors.

volatility measure. For Panel B, and Panel C, we use respectively LVUS and LVDM as volatility measures. To give economic meanings to the estimated vectors, they are normalized on exports, X_t. This is done by setting the estimated coefficient on X_t by -1 and dividing each cointegrating vector by the negative of the estimated X_t coefficient.

Parameter estimates that represent long run elasticities, together with their respective *t*-values are shown in Table 3. The estimated income elasticity for all measures of volatility is positively related to export volume. The long run income elasticity is significant and greater than unity, and implies a large response of total exports to changes in foreign income.

Our estimate of relative price is positive and is not significant at the 5% level when we use LV as a measure of volatility. In the case of nominal volatility, our estimates of relative price are negative but not significant at the 5% level. As noted in Arize *et al*. (1997), the insignificant price effects are generally attributed to at least three factors. The first is the use of unit-value indexes, which are computed from observation units where some aggregation has taken place. They are

accurate only if the composition of the unit remains the same or if the net effect of such changes is insignificant. The second is that as a developing country, Turkey may have been able to differentiate her exports by focusing on nonprice factors such as one-time delivery, design improvement, product varieties, and aggressive marketing. Finally, price elasticity that is positive and/or insignificant can certainly be the result of poor data quality.

Table 3
Estimate of the Cointegrating Relationship

	Normalized Cointegrating Vector
A:	$\ln X = 4.926\ln Y + 1.802\ln P + 0.433\ln V$ (15.69) (1.66) (4.56)
B:	$\ln X = 2.733\ln Y - 3.062\ln P + 0.297\ln VUS$ (3.07) (0.96) (2.87)
C:	$\ln X = 2.601\ln Y - 2.149\ln P + 0.329\ln VDM$ (2.08) (1.21) (2.81)

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

An important aspect of the results is that the elasticity estimate of the exchange-rate volatility has positive sign and is statistically significant at each measure of volatility. This result suggest that 43.3% (29.7% and 32.9%) of all Turkish export is affected by real (nominal) exchange rate volatility. This would seem to indicate that export firms based in Turkey have responded to exchange-rate volatility by increasing export. Overall, the results further indicate that exports are more responsive to the change in real effective exchange rates with the currencies of its major trading partners[15].

The positive impact of volatility on export provides support for the hypothesis that volatility is not treated simply as a trading risk by most Turkish exporters. One possible explanation of the positive volatility impact on exports would be that producers supplying export markets in Turkey are aware that they cannot rely on domestic market to absorb any excess supply that might occur if trading becomes more risky due to increased exchange rate volatility. Hence, to avoid any reduction in revenues arising from increased risk, they may export more.

[15] Özbay (1999) investigates possible effects of exchange rate uncertainty on exports for Turkey in the context of the GARCH model for the period 1988-1997. Her findings indicate that exports are adversely affected by real exchange rate uncertainty.

3.2. Error correction model (ECM)

Using the cointegrating vector normalized on exports, we estimated an error correction model (ECM) that provides us with information on the short run export function. The results are summarized in Table 4. Based on the Representation Theorem developed in Engle and Granger (1987), it can be shown that, if a cointegrating relationship exists among a set of I(1) series, then a dynamic ECM representation of the data also exists. Following Hendry's (1995) general-to-specific modeling approach, we first include four lags of the first-difference of each variable in equation (1), a constant term and lagged error-correction term (EC_{t-5}) generated from the Johansen procedure, and then gradually eliminate the insignificant variables[16]. This allowed us to derive a parsimonious model. Before we discuss the results, we need to determine the adequacy of the ECM. Several diagnostic tests were performed and reported in the last column of Table 4. Diagnostic tests indicate that the model is correctly specified. The adjusted R^2 for LV, LVUS, and LVDM are 0.66, 0.74 and 0.73, respectively. These values compare well with those reported in other studies for regressions based on first differences in variables (see for example Arize, 1995; Doyle, 2001; and Bredin *et al*., 2003).

Given the evidence supporting the adequacy of the estimated ECM, we can make a number of observations regarding the estimates in Table 4. First, the coefficient of the error correction term is significant and has the correct sign for all volatility measures. The significance of the lagged error correction term displaying the appropriate negative sign supports the cointegration findings and implies a valid equilibrium relationship between the variables in the cointegrating equations. This means that excluding the cointegration relationship would have led to misspecification in the dynamic structure of the model. The coefficient on the error correction term indicates what proportion of the discrepancy between the actual and long run or equilibrium value of exports is eliminated or corrected each quarter. The result indicates that the adjustment of export volume to changes in the regressors may take about three quarters in Turkey when we use the real exchange rate volatility measure. The result also indicates the existence of market forces in the export market that operate to restore long run equilibrium after a short run disturbance.

[16] Since we used four lags to determine the number of cointegrating vectors, then error correction term should be $EC_{t-(n+1)} = EC_{t-5}$, where n denotes the number of lags.

Table 4
Estimated Error-Correction Model

lag	EC (-1)	Δ ln *X*	Δ ln *Y*	Δ ln *P*	Δ ln *V*	Summary Statistics
Volatility measure: LV						
1	-0.292(-2.832)		1.550 (1.39)			$\overline{R}^2$ =0.66
2		-0.212 (-2.56)				ARCH(4)=3.89(0.77)
3		-0.230 (-2.75)			0.044 (1.90)	DW=2.005
4		-0.402 (-4.56)		-0.859(-2.54)		RESET=1.29 (0.36)
						NORM=0.164 (0.33)
Volatility measure: LVUS						
1	-0.429 (-4.40)		2.371 (2.05)		-0.074(-1.37)	$\overline{R}^2$ =0.74
2				0.575 (1.91)		ARCH(4)=1.93(0.75)
3		-0.204 (-2.95)				DW=2.016
4		0.458 (5.87)		-0.740(-2.46)		RESET=3.44 (0.48)
						NORM=1.017 (0.60)
Volatility measure: LVDM						
1	-0.482(-4.92)					$\overline{R}^2$ =0.73
2				0.535(3.11)	-0.107(-1.51)	ARCH(4)=1.64(0.18)
3		-0.205(-3.08)				DW=2.083
4		0.407(5.62)		-0.778(-2.60)		RESET=1.44 (0.24)
						NORM=1.544 (0.46)

Note: Figures in parentheses are the t-statistics. The critical value at 10% is 1.29 and 1.66 at 5% (1-tail).

Second, the results indicate that changes in foreign income have positive and statistically insignificant short run effects on exports when we use LV. However, when we use LVUS as a measure of exchange rate volatility, the results suggest a positive and significant relationship between exports and foreign income. The results further indicate that foreign income has no short run effects on exports when we use LVDM as a measure of volatility. Third, in contrast to the long run results, changes in relative prices have negative and statistically significant short run effects on exports in all measures of volatility. Finally, the real exchange rate volatility appears as a statistically significant variable. We conclude that volatility has a positive short run effect on exports. Volatility supports its long run effect established earlier when we use LV and has insignificant negative signs for LVUS and LVDM. The positive relationship between real exchange rate volatility and exports requires exports to increase (decrease) as the exchange rate depreciates

(appreciates). This result confirms earlier findings reported in Doyle (2001). She examines the impact of exchange rate volatility on Ireland's exports to the United Kingdom and finds that both real and nominal volatility were important determinants of Irish-UK trade, with positive effects predominating. Bredin *et al*. (2003) also examine the long run and short run relationship between export volume and exchange rate volatility in Ireland for the period 1979-1992. They find that real exchange rate volatility has no effect on the volume of trade in the short run but a significant positive effect in the long run. Our findings also have theoretical support. Franke (1991), Sercu and Vanhulle (1989) and Dellas and Zilberfarb (1993) have developed models showing how increased exchange rate volatility may actually lead to greater levels of trade.

4. Conclusions

This paper investigated the impact of exchange rate volatility on Turkey's exports to its major trading partners for the period 1982-2001, using the techniques of cointegration and error correction methods. The volatility term is defined as the moving standard deviation of the growth of the real and nominal exchange rates. The estimated cointegration vectors imply that there exists a unique long run relationship between the volume of exports, income of the foreign countries, relative prices and exchange rate volatility. The volatility coefficients have positive signs across all measures. Among the three measures of volatility, the results further indicate that exports are more responsive to the change in real effective exchange rates with the currencies of its major trading partners. The negative signs of the error correction terms indicate a significant negative effect on export demand in the short run. The change in exports volume per quarter that is attributed to the disequilibrium between the actual and equilibrium levels is measured by the absolute values of the error correction term of each equation. The results suggest that the adjustment of export volume to changes in the regressors may take about three quarters when we use LV (real exchange volatility) as a measure of volatility. The results point to the existence of market forces in the export market that operate to restore the long run equilibrium after a short run disturbance.

Our findings further suggest that there is a positive and statistically significant short run relationship between exports and exchange rate volatility. The findings of a long run and short run positive impact of real exchange volatility on exports are consistent with the results reported in Asseery and Peel (1991), Doyle (2001) and Bredin *et al*. (2003).

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Appendix

Calculating the real effective exchange rate for Turkey

1. E denotes the bilateral exchange rates (E) between Turkey's (country j) currency and the currencies of its major trading partners (country i's).

2. We calculate the real bilateral exchange rates RE_{ij} using E_{ij}'s and CPI indexes: $RE_{ij} = [(E_{ij} \cdot CPI_i) / CPI_j]$.

3. We construct the index of real bilateral exchange rates by selecting 1995 as the base period: $REER_{ij} = \left(RE_{ij}^{t} / RE_{ij}^{95}\right) \cdot 100$.

4. We take the weighted average of $REER_{ij}$ in order to obtain the index of real effective exchange rates for Turkey:

$$REER_j = \sum_{i=1}^{n} \delta_{ij} REER_{ij}$$ where δ_{ij} are the weights.

Note that an appreciation of the real effective exchange rate is reflected by a decrease of the index ($REER_j$) and depreciation by an increase of the index ($REER_j$).

Özet

Türkiye'de döviz kuru belirsizliği ve ihracat hacmine etkisi

Bu makale Türkiye'nin en önemli ticaret ortaklarına yaptığı ihracat üzerinde reel döviz kuru oynaklığının etkisini 1982-2001 dönemine ait çeyreklik veriler kullanarak incelemektedir. Eşbütünleşme ve hata düzeltme modellemesi yaklaşımları uygulanmıştır. Eşbütünleşen (eşbütünleşik) ilişkiler Johansen'in çok değişkenli yöntemi kullanılarak tahmin edilmiştir. Kısa dönemli dinamikler hata düzeltme tekniklerinden yararlanılarak tahmin edilmiştir. Bulgularımız, döviz kuru oynaklığının ihracat hacmi üzerinde uzun dönemde anlamlı pozitif etkisi olduğunu göstermektedir. Bu sonuç, Türkiye gibi küçük bir ekonomide faaliyet gösteren firmaların artan döviz kuru riskiyle baş etmek için fazla seçeneğinin olmadığını göstermektedir.