
THE IMPACT OF EXCHANGE RATE VOLATILITY ON EXPORT FLOWS IN DEVELOPING COUNTRIES

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Abstract

This article investigates the long run impacts of exchange rate volatilities on export volumes in six developing countries, Turkey, Indonesia, Brazil, Mexico, Poland, and Chile. The exchange rate volatility is estimated by both ARCH/GARCH and moving sample standard deviation-based volatility proxies. Error correction representation of the long run export model reveals that there is not cointegration relationship for Brazil, Indonesia, and Chile while the results support the cointegration relationship for Turkey, Mexico, and Poland. Second, world income is major determinant driving export volumes in Turkey, Mexico, and Poland. Third, there is a significant negative relationship between exchange rate level and export volumes in Turkey and Mexico in line with theory. Finally, the measure of exchange rate volatility carries a significant negative coefficient in the case of Poland but a significant positive coefficient in the case of Turkey.

Jel Codes: F14, F31

Keywords: Real exchange rate, exchange rate volatility, export, long run equilibrium

DÖVİZ KURU OYNAKLIĞININ GELİŞMEKTE OLAN ÜLKE İHRACAT AKIMLARINA ETKİSİ

Öz

Bu çalışma döviz kuru oynaklığının ihracat akımları üzerindeki uzun dönem etkisini Türkiye, Endonezya, Brezilya, Meksika, Polonya ve Şili geliştirmekte olan ülkeleri özelinde incelemektedir. Döviz kuru oynaklığı, hem ARCH/GARCH modeli hem de hareketli standart sapmalar bazlı ölçümlerle elde edilmiştir. Hata düzeltme modeli sonuçları Brezilya, Endonezya ve Şili özelinde bir eşbütünleşme ilişkisinin varlığını tespit edemezken diğer ülkelerde eşbütünleşme ilişkisi bulunmuştur. İkinci olarak, dış dünya gelir düzeyi Türkiye, Meksika ve Polonya'da ihracat akımlarını etkileyen en etkili değişkendir. Üçüncü olarak, Türkiye ve Meksika'da teorik beklentilere uygun şekilde reel döviz kuru düzeyi ile ihracat akımları arasında anlamlı negatif ilişkiler bulunmuştur. Son olarak ise, döviz kuru oynaklığı Polonya'da ihracat akımlarını olumsuz etkilerken, Türkiye'de ihracat üzerinde anlamlı pozitif ilişkiler üretmektedir.

JEL Sınıflandırması: F14, F31

Anahtar Kelimeler: Reel Döviz Kuru, döviz kuru oynaklığı, ihracat, uzun dönem denge

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1. Introduction

Since the beginning of floating exchange rate regimes due to the collapse of the Bretton Woods System, exchange rates in many economies have started to fluctuate freely. Since then, uncertainty of the exchange rates and their impacts on different macroeconomic variables have emerged as one of the major research areas in economics. For instance, many empirical and theoretical studies have been published to reveal the link between “volatility of exchange rates” and “foreign trade volumes”. Theoretically, the direct effect remained ambiguous, and direction of the impact was explained by the size of the “substitution and income effects” of the volatilities. This suggests that the direction of the link between volatility of exchange rate and trade volumes can only be resolved empirically. According to the large number of empirical studies, no general consensus has been reached on this matter, that is, volatility of exchange rate may produce either positive or negative impacts on trades. Investigation of the direction of the link is crucial for economies since policy-makers can derive different policy implications based on the country-specific evidences. For example, as pointed out by Bahmani-Oskooee and Gelan (2018), if volatility of exchange rate has an adverse effect on trade, a Central Bank often intervenes foreign exchange market to dampen the volatility and aims to increase trade flows. However, if exchange rate volatility has positive impacts on trade flows, then completely free exchange rates and no foreign market intervention will be quite efficient. As a result, what I do in this paper is to empirically reveal the link between volatility of exchange rates and exports.

This article investigates the long run impacts of exchange rate volatilities on export volumes for six developing countries, Turkey, Indonesia, Brazil, Mexico, Poland, and Chile² using quarterly data by conducting autoregressive distributed lag (ARDL) model and error correction representation (ECM). The volatilities of many financial times series including exchange rates are not constant over time, and ARCH/GARCH models allows volatility to vary through time. However, when exchange rate data do not provide ARCH effects, the exchange rate volatility measure will be substituted by another volatility measure, moving sample standard deviation. Hence, the main contribution of this study is to employ both ARCH/GARCH and moving sample standard deviation-based measures together when modeling directly unobservable exchange rate volatilities in developing countries. Finally, long run estimates of the export function are produced from ARDL-bound testing approach and error correction representation based on these two volatility proxies.

The remainder of the chapter is as follows. Section 2 explains the theoretical link between exchange rate volatility and export volumes. In Section 3, I present a brief overview of the empirical literature. Section 4 includes model specification and selected econometric method while data and empirical results are reported in Section 5. Finally, summary and conclusion are drawn in Section 6.

2. Theory: Effects of Real Exchange Rate Volatility on Export Volumes

In general, the theoretical link between volatility of exchange rates and export volumes can be explained by size of the “income and substitution effects”. Regarding the impacts of exchange rate volatility and trade volumes, one of the leading theoretical analyses has been outlined by Hooper and Kohlhagen (1978) under the theory of uncertainty. They argue that if exporters are risk averse then elevated volatility in exchange rate is expected to reduce export flows. As explained by Arize, Osang, and Slottje (2000), exchange rate is determined by the exporter and importer at the time of trade contract. However, exporters do not receive their payments until the exported goods and services are delivered. Under the existence of high volatility, unpredictable exchange rates produce uncertainty about the future earnings and discourage exporters from selling their products abroad. This negative

² Detailed information on the classification of developing countries can be found in the International Monetary Fund's World Economic Outlook (2018, pp. 134-135).

impact on trade volumes explains the “substitution effect” of volatility of exchange rates on trade. However, Ethier (1973) and Baron (1976) argue that if the forward markets are available and hedging is possible, exchange rate uncertainty does not produce any impact on trade flows. Viaene and DeVries (1992), on the other hand, point out that if the hedging is difficult and expensive, uncertainty has indirect impact on trade volumes, and substitution effect is still valid.

Another possible impact of the volatility of exchange rate on trade volumes is explained by “income effect”. According to the income effect, if exporters are risk averse, a higher volatility elevates the expected marginal utility of export revenue as a result of quick and massive decline of export volumes driven by substitution effect. Then, increasing marginal utility of export revenue stimulates exporters to increase their products that they sell abroad. This effect is known as “income effect” of the volatility of exchange rates on export flows. Consequently, the net effect is determined by the sum of the magnitude of the income and substitution effects. Based on the net impact, exchange rate volatility may produce either a positive or a negative impact on export volumes.

3. Literature

One of the very first paper to explore the relationship between exchange rate volatility and trade is published by Hooper and Kohlhagen (1978). They analyze the exchange rate uncertainty on prices and trade between U.S. and Germany over the period of 1965–1975. Hooper and Kohlhagen conclude that exchange rate uncertainty produces significant impact on prices while results does not confirm the link between uncertainty and trade volumes. Chowdhury (1993) studies the impact of exchange rate volatility on the trade flows of the G-7 countries. He reveals the negative impacts of exchange rate volatility on export volumes in each of the seven countries. He also asserts that previous studies consistently find quite weak relationship between trade volumes and volatility since these studies do not take into account the stochastic properties of the relevant time series. Franke (1991) argues that elevated exchange rate volatility might increase the trade level of exporting firms. He also presents some conditions that exchange rate volatility stimulates the volume of international trade. Similarly, De Grauwe (1988), Dellas and Zilberfarb (1993) support the argument that volatility may increase the volume of trade while Gagnon (1993) and Wolf (1995) suggest that volatility dampens the trade volumes.

Arize, Osang, and Slottje (2000) use Johansen’s cointegration approach for 13 less developed countries, and their results suggest that volatility significantly reduces the export volumes in each of the 13 countries. Doroodian (1999) studies the influences of exchange rate volatility on trade volumes for developing countries, over the period of 1973 and 1996 by employing ARMA and GARCH models. He concludes that uncertainty on exchange rate significantly reduce trade volumes. Doganlar (2002) investigates the impacts of exchange rate volatility for Turkey, South Korea, Malaysia, Indonesia and Pakistan by using Engle-Granger cointegration approach. The result presents that exchange rate volatility significantly reduces the export volumes. Wang and Barrett (2007) investigate the impact of exchange rate volatility on Taiwan’s exports to the U.S. between 1989-1998 by using multivariate GARCH-M estimator. They find that change in either importing country production or expected exchange rate are the main determinants of trade flows. Additionally, they claim that exchange rate volatility only influences the agricultural trade volumes. Asteriou, Masatci and Pilbeam (2016) employ GARCH-based volatility measure in the ARDL error correction representation for Mexico, Indonesia, Nigeria, and Turkey. They find no relationship between volatility and foreign trade except for Turkey in the long run. Arize, Osang, and Slottje (2008) use several cointegration techniques over the period of 1973-2004 for the eight Latin American countries, and suggest that increase in exchange rate volatility significantly decreases export flows in the long run. Hall, Hondroyannis, Swamy, Tavlas and Ulan (2010) show that there is not a negative and significant relationship between Emerging Market Economies

exchange-rate volatility and exports flows for the period between 1980 and 2006. Additionally, they conclude that open capital markets might be a major factor reducing the negative impact of volatility on trades.

Houchet-Bourdon and Korinek (2011) investigate the relationship between volatility and trade flows in China, the Euro area and the United States by using ARDL bound test and error correction representation. They find that volatility weakly affects trade. Additionally, the impact of exchange rates on exports of agricultural goods relatively greater than manufacturing goods. Finally, exchange rate volatility does not present a strong evidence to stimulate bilateral trade flows. Bahmani –Oskooee and Gelan (2018) study the exchange rate volatility and trade relationship for the twelve African countries. They find evidence that volatility produces significantly negative impact on export demand in two countries and positive impacts on three countries. They also state that major component which triggers export volumes is world income. Baum and Caglayan (2010) analyze the industrialized countries for the time period 1980-2006, and conclude that volatility of exchange rate produces a significantly positive impact on the volatility of bilateral trade volumes. However, exchange rate volatility does not exert a significant effect on the level of the trade volumes.

4. Model Specification and Econometric Method

With regard to modelling the impact of the level of exchange rates and their volatility on export demands in developing countries, I follow a traditional long-run equilibrium export model proposed by Arize (1995) and Chowdhury (1993). This model depends upon the flexible exchange-rates both in terms of level and volatility and world income. Long run export volume are formulated as follows:

$$\log X_t = b_0 + b_1 \log W_t + b_2 \log RER_t + b_3 \text{VOL}_t + \omega_t \quad (1)$$

where X_t is a country's real exports, W_t is real world income, RER_t is relative prices, i.e. real effective exchange-rate, and VOL_t is the exchange rate volatility. ω_t is, on the other hand, a disturbance term. If the real world income raises, it should stimulate more exports; hence, its sign is expected to become positive, i.e. $b_1 > 0$. As a decrease in exchange rate represents a real depreciation of home currency, RER_t variable should carry negative sign, i.e. $b_2 < 0$. Finally, since the impacts of exchange rate volatility on export is uncertain in literature, an estimate of b_3 can reveal either a negative or positive sign. Finally, all variables in equation (1) are expressed in logarithmic form.

In this study, I follow the ARDL-bound testing approach proposed by Pesaran and Shin (1998) and Pesaran, Shin, and Smith (2001) and error correction representation. Equation (1) is reformulated as the ECM form in equation (2):

$$\Delta X_t = d_0 + \alpha \text{ECM}_{t-1} + \sum_{i=1}^p d_i^X \Delta X_{t-i} + \sum_{i=0}^{q_1} d_i^W \Delta W_{t-i} + \sum_{i=0}^{q_2} d_i^{\text{RER}} \Delta RER_{t-i} + \sum_{i=0}^{q_3} d_i^{\text{VOL}} \Delta \text{VOL}_{t-i} + u_t \quad (2)$$

$$\text{ECM}_{t-1} = \theta_1 X_{t-1} - \theta_2 W_{t-1} - \theta_3 RER_{t-1} - \theta_4 \text{VOL}_{t-1}$$

where α is speed of adjustment coefficient. The lagged level variables θ_2 , θ_3 , and θ_4 are normalized on θ_1 , and determine the long run effects the export model³. The optimal lag orders p , q_1 , q_2 and q_3 can be specified by the Schwarz-Bayesian (1978) (SBIC) or the Akaike (1973) information criterions (AIC). According to the ECM, if the current amount of the export is not equal to the fundamental value or

³ Error correction representation in (2) also gives short run effects of exogeneous variables on exports, i.e. d_i^W , d_i^{RER} , d_i^{VOL} ; however, in many studies, either the results are statistically insignificant or estimated statistically significant coefficients carry an incorrect sign. As a result, many studies conclude that short-run effects do not track any pattern (Baek and Koo (2009), Bahmani-Oskooee and Gelan (2018), Huchet-Bourdon and Korinek (2011), Bahmani-Oskooee and Ratha (2004), Simsek and Kadilar (2006). Additionally, exporters can eliminate the exchange rate risk in short run easily as hedging in short run is not costly relative to long run (Perée and Steinherr (1989)). Consequently, the literature pays much more attention to the long run impacts of exchange rate volatility on trade volumes.

long run amount of the export volume, then current export will adjust itself toward long run equilibrium level by the rate of speed of adjustment coefficient. Therefore, any deviation from the long run will correct itself through time.

ARDL model is being used for decades, and has some practical advantages for the researchers. First, ARDL model has a reparameterization in the form of error correction representation, and the presence of cointegrating relationship can be analyzed based on the error correction form. Second, ARDL model does not necessitate to identify that whether the variables of interest are integrated of order zero or one, i.e. I(0) or I(1), respectively unlike the conventional cointegration methods⁴. Hence, I(0) or I(1) variables can be employed in a same estimation. Third, unlike the conventional methods such as Johansen (1988) maximum likelihood which determines the cointegration relationship in a system equations, ARDL method uses a single equation time series form. Fourth, ARDL model produces more efficient estimators when the sample length is not very large. Finally, the inclusion of short term dynamics in the form of lagged-differenced series in the ARDL model may handle possible endogeneity and serial correlation problems⁵.

To verify the existence of a cointegration relationship, the conventional F test for joint significance of lagged-level variables is conducted. The equation (2) is first estimated by using ordinary least square (OLS), and the null hypothesis of no cointegration, i.e. $H_0: \theta_1 = \theta_2 = \theta_3 = \theta_4 = 0$, is tested by using F test against alternative hypothesis, i.e. $H_0: \theta_1 \neq \theta_2 \neq \theta_3 \neq \theta_4 \neq 0$. Since the conventional F test is inapplicable in the context of ARDL approach, Pesaran et al. (2001) tabulates two distinct critical values which are called upper and lower bound limits. According to the ARDL-bound testing approach, cointegration exists if the computed F statistic exceeds the upper bound limit. If the calculated F statistic less than the lower bound limit, then one cannot reject the the null hypothesis of no cointegration. Finally, if the calculated F statistic takes a value inside the upper and lower bounds, then the test does not conclude.

After detecting the existence of a cointegrating relationship in the ECM form of the export model (2), the long run coefficients of the model can be estimated. A negative and significant speed of adjustment coefficient (α) confirms the equilibrium relationship and present another hint for the existence of cointegration relationship among variables. Finally, the post estimation diagnostics are generated to establish how the export model (2) fits the data. These diagnostic tests are Ramsey Reset Test for confirming the linear specification of the model against nonlinear specification proposed by Ramsey (1969), and Breusch–Godfrey Lagrange Multiplier Test for testing serial correlation free residuals, and White (1980) test for heteroscedasticity.

4.1. Measurement of Exchange Rate Volatility

Although there are large number of studies investigating the impact of exchange rate volatility on various macroeconomic variables, there is still no consensus on what is the best measure of “volatility”⁶. In this article, I use two broadly used distinct proxies to measure exchange-rate volatility.

⁴ Engle and Granger (1987) two-step process and Johansen (1988) maximum likelihood approach are known as the conventional cointegration methods in literature and both require to determine the integrated of orders of the variables before the estimation process. The standard statistical inference based on these methods are depends on the argument that all variables are integrated of order one. Additionally, general pretesting problems such as misclassification of variables as I(0) or I(1) may lead researchers to conclude incorrect statements about time series feature of the data (Kripfganz and Schneider (2016)).

⁵ In the error correction model (2), several instrumental variables proposed by Pagan and Ullah (1988) are used under the assumption of endogeneity of some of the explanatory variables. These instrumental variables are constant coefficient, the lagged error correction term (ECM_{t-1}), and explanatory variables with a number of lagged-differences.

⁶ Some of the example of these distinct measures are “standard deviation of the first difference of the log exchange rates”, “moving sample standard deviation measure of the log exchange rates”, and “conditional volatility measure estimated from

First measure is known as the conditional volatility of exchange rates estimated from a generalized autoregressive conditional heteroscedasticity (GARCH) model proposed by Bollerslev (1986). Time-dependent volatility of the exchange rate as a function of observed prior volatility is broadly estimated by ARCH models proposed by Engel (1982). Later, Bollerslev extended ARCH models to a GARCH model by adding lagged values of the conditional variance. GARCH models and their extensions were empirically considered quite successful in modelling volatility clustering feature of the exchange rate data. Volatility clustering is first stated by Mandelbrot (1963), and it represents the fact that large changes are tend to be followed by large changes and small changes are tend to be followed by a small change⁷. A general GARCH(m; k) process for exchange rate is written as:

$$\Delta RER_t = \omega + \varepsilon_t \quad (3)$$

$$\sigma_t^2 = \gamma_c + \gamma_1 \varepsilon_{t-1}^2 + \gamma_2 \varepsilon_{t-2}^2 + \dots + \gamma_m \varepsilon_{t-m}^2 + \zeta_1 \sigma_{t-1}^2 + \zeta_2 \sigma_{t-2}^2 + \dots + \zeta_m \sigma_{t-k}^2 \quad (4)$$

$$\varepsilon_t \sim \text{GED}(0, \sigma_t^2, \rho)$$

Where ε_t is disturbance term, ω is a constant and γ_i are the ARCH parameters and ζ_i are the GARCH parameters. GARCH(m; k) model contains the combination of conditional mean equation (3) and conditional variance equation (4). As seen in equation (4), m-lagged values of squared disturbances and k-lagged values of conditional variances form the current conditional variance (σ_t^2) of the regression (3) disturbances. The conditional distribution of the disturbance term has a vital role when estimating the GARCH models. When estimating the model, I employ Generalized Error Distribution (GED) Distribution with shape parameter ρ since the distribution of the most of the financial time series in practice are leptokurtic, i.e. extreme values are more frequent than normally distributed series. The GED distribution with shape parameter smaller than two has fatter tails; hence, it asymptotically converges to zero more slowly than Gaussian normal⁸. In this study, conditional volatility is measured by equation (4). Finally, the optimal lags of equation (4) is determined by likelihood ratio test.

Second measure, on the other hand, is known as moving sample standard deviation (MSSD) of volatility expressed in the form of:

$$\text{VOL}_t = [(1/p) \sum_{i=1}^k (\text{RER}_{t+i-1} - \text{RER}_{t+i-2})^2]^{1/2} \quad (5)$$

where p is the order of moving average⁹. Whenever exchange rate data do not exhibit ARCH effects then I use conditional volatility equation (5) instead of (4).

5. Data and Empirical Results

5.1. Data

The data used are quarterly for the sample period 1998:Q1-2017:Q1 for Turkey; 1996:Q1-2017:Q1 for Mexico; 2000:Q1-2017:Q1 for Indonesia; 1996:Q1-2017:Q1 for Brazil; 2002:Q1-2017:Q1 for Poland;

ARCH/GARCH models". For a detailed information for distinct measures, see Mckenzie (1998), Huchet-Bourdon and Korinek (2011), Arize et al (2000), Nicita (2013), Sauer and Bohara (2001), Hall et al.(2010)

⁷ For example, Pozo (1992), Dorodian (1999), Wang and Barrett (2007), Rahmatsyah, Rajaguru, and Siregar (2002), Hall et al. (2010), Kroner and Lastrapess (1993), McKenzie and Brooks (1997), Mckenzie (1998) have used ARCH/GARCH-based measure of volatility.

⁸ Other distributions having fatter tails than Gaussian normal distribution are Laplace and Student-t distributions.

⁹ I employ the eight-quarter moving standard deviation of the log real effective exchange rate in this study. This measure is quite standard in much of the literature (e.g. Arize et al (2000), Sauer and Bohara (2001), Arize, Malindretos, and Kasibhatla (2003), Rey (2006), Hall et al. (2010), Bahmani-Oskooee and Gelan (2018), Chowdhury (1993), Kenan and Rodrik (1986), Doganlar (2002)).

and 2000:Q1-2017:Q1 for Chile based on the data availability. All data are represented in real terms and obtained from Federal Reserve Bank of St. Louis and the Central Bank of the Republic of Turkey¹⁰.

5.2. Empirical Results

After differencing the log exchange rate to correct for nonstationarity¹¹, one can present the descriptive summary statistics for the log-differenced exchange rate series. Table 1 shows that all series exhibit the usual features of GARCH models such as excess kurtosis (or leptokurtic distribution) and negative skewness¹². While Shapiro-Wilk (1965) W test for normality reveals that log-differenced series are not normally distributed at ten percent significance level each of the six countries. Moreover, Engle's (1982) Lagrange multiplier test (ARCH-LM Test) results confirm the presence of autoregressive conditional heteroskedasticity in the results for Turkey, Indonesia and Poland at ten percent significance level. As stated by Westerfield (1977) and Hsieh (1989), leptokurtic exchange rate differences suggest volatility clustering, and GARCH-based models are the quite convenient for modelling the volatility clustering feature of the financial time series. As a result, all these preliminary results suggest that GARCH models are best fit for the exchange rate volatility for Turkey, Indonesia and Poland. On the other hand, I choose the moving sample standard deviation as a measure of volatility of exchange rate over the GARCH-based volatilities for Brazil, Mexico and Chile data since ARCH-LM tests for these countries do not exhibit ARCH effects on the log-differenced exchange rate data. Now, one can further estimate the conditional variance equation (4) for Turkey, Indonesia and Poland.

Table 1: Descriptive Statistics on Log-Differenced RER

	Turkey	Indonesia	Brazil	Mexico	Poland	Chile
Skewness	-0.232	-1.979	-1.81	-0.709	-1.375	-0.482
Kurtosis	4.432	11.66	9.227	4.77	6.306	3.504
Shapiro-Wilk W Test statistics^a	0.97***	0.823*	0.866*	0.960**	0.896*	0.977***
ARCH-LM Test Statistics^b	6.568*	3.333***	0.07	0.990	10.394*	0.589

Note: ***, Significance at 10%; **, significance at 5%; *, significance at 1%.

^a Null hypothesis Shapiro-Wilk (1965) W Test is "data are normally distributed".

^b Null hypothesis for the Engle's (1982) ARCH-LM test is "No ARCH effect".

The model of choice to measure exchange rate volatility in this study is the ARCH(1) for Turkey; GARCH(1,1) for Indonesia; and GARCH(1,2) for Poland based on the likelihood ratio test¹³. Table 2 summarizes the results obtained from conditional variance equation (4) and presents several

¹⁰ For detailed data definitions and data sources, see Appendix A, Table A.1.

¹¹ Augmented Dickey Fuller (1979) (ADF) and Phillips-Perron (1988) unit root test results for log real effective exchange rate are reported in Appendix A. The results are given based on three distinct cases: trend and constant, constant, and none.

¹² A distribution with excess kurtosis implies that tails are fatter than normal distribution; therefore, such a distribution produces more outliers than the normal distribution. If a calculated kurtosis for a distribution is above 3, then the distribution is said to be "leptokurtic". Negative skewness implies that a distribution has more extreme measurements in the left tail than right tail.

¹³ One can argue that the sample periods are not large enough for ARCH/GARCH volatility estimates. However, several studies in the literature employ similar sample sizes when estimating exchange rate volatilities. For example, Arize (1995), Hall et al. (2010), Doroodian (1999), Crowley and Lee (2003) and Rahmatyash et al. (2002) use quarterly data for the sample lengths varying between 18 and 26 years

regression diagnostics. λ represents the number of iterations that log-likelihood convergence is achieved while $\ln L$ denotes the value of the log-likelihood function. All three estimated shape parameters are less than two which implies that the distributions of the disturbances have fatter tails than normally distributed disturbances, as expected. This also indicates that the choice of GARCH with GED distribution over the GARCH with Gaussian normal appears to be quite reasonable. Finally, ARCH-LM test reveals that one cannot reject the null of no ARCH effects in the standardized residuals, and the data for Turkey, Indonesia, and Poland fit the model very well.

Table 2: GARCH Estimations, Conditional Volatility of RER

	Turkey	Indonesia	Poland
ARCH(1)	0.593 ***	1.405*	0.593**
ARCH(2)			
GARCH(1)		0.159***	0.750*
GARCH(2)			-0.273*
λ	9	39	18
$\ln L$	173.766	146.99	172.604
ρ	1.115	1.750	1.859
ARCH LM Test Stat.	0.019	1.492	2.123

Note: ***, Significance at 10%; **, significance at 5%; *, significance at 1%.

Table 3: ARDL Bound Test and Several Diagnostics

	Turkey	Mexico	Indonesia	Brazil	Poland	Chile
Volatility Measure	ARCH(1)	MSSD	GARCH(1,1)	MSSD	GARCH(1,2)	MSSD
Speed of Adjustment Coefficient (α)	-0.690*	-0.259*	-0.307*	-0.097	-0.433*	-0.050
F Statistics ^a	12.327*	9.607*	3.270	2.070	9.656*	1.698
Breusch-Godfrey LM Test Statistic ^b	$l=1$ 0.006	0.014	1.038	3.180	2.865	0.790
Ramsey's Reset Test Statistic ^c	$l=4$ 2.001	1.291	5.610	3.852	3.495	4.543
White Test Statistics ^d	1.30	0.16	2.01	1.46	1.69	0.26
	69.20	66.85	16.03	46.68	1.61	33.84

Note: ***, Significance at 10%; **, significance at 5%; *, significance at 1%. Optimal lag order for the error correction model is selected by SBIC, and the selected lag orders are ARDL(4,1,1,0) for Turkey, ARDL(1,2,2,0) for Mexico, ARDL(1,1,0,0) for Indonesia, ARDL(2,1,4,0) for Brazil, ARDL(1,0,2,0) for Poland, and ARDL(1,1,1,0) for Chile.

^a Lower and critical upper critical values for F test are 5.17 and 6.36 at one percent significance level, 4.01 and 5.07 at five percent significance level, and 3.47 and 4.45 at ten percent significance level respectively. All models include unrestricted constant and trend. Null hypothesis indicates no cointegration relationship.

^b Null hypothesis of the Breusch-Godfrey LM test is “no serial correlation”. Test statistics are given for two different lags, $l=1$ and $l=4$.

^c Null hypothesis of the Ramsey’s (1969) Reset test is “No functional form specification”.

^d Null hypothesis of the White’s (1980) Test for heteroscedasticity is “residuals have constant variance.”

Given the presence of the real effective exchange rate volatilities obtained from either GARCH or MSSD based calculations in the error correction model (2), Table 3 presents the ARDL-bound testing results and several diagnostics. When estimating an ARDL model, all statistical results are sensitive to violation of classical assumptions on disturbance term. As seen in Table 3, bound test results justify the cointegration relationship in the results for Turkey, Mexico, and Poland since F-values are greater than the upper critical value. Additionally, both significant and negative α coefficients for these countries indicate the adjustment toward long run equilibrium and presents another evidence that cointegration relationships are valid. If a shock appears in short run, the estimated α coefficient obtained from Turkey’s export model concludes that about 70 percent of the deviation from the long run equilibrium corrects itself in one quarter. This rate is 26% and 43% for Indonesia and Poland, respectively. Finally, ARDL-bound testing results, on the other hand, do not confirm the validity of cointegration relationship for Indonesia, Brazil, and Chile since F-values fall below the lower bounds. To confirm whether or not the findings obtained from error correction form of the export model in Table (3) are valid, one should conduct diagnostic tests. First, Breusch-Pagan test result exhibits that null hypothesis of no serial correlation cannot be rejected for each of the six countries, and residuals are free from serial correlation. Second, Ramsey’s Reset Test (1969) reveals that F statistic is lower than the critical value indicating that there is not a functional form misspecification, i.e. non-linear combinations of the explanatory variables cannot explain the export variable. Finally, White (1980) test results shows that one cannot reject the null of constant variances for each of the six countries. Consequently, all diagnostic tests provide support for the ARDL error correction representation, and the statistical fit of the model to the data is quite reasonable.

Table 4: Long-Run ECM Estimates

Long Run Estimates (P-value)	Turkey	Mexico	Poland
W	2.874*	0.800***	1.07**
RER	-0.454*	-0.633*	0.097
VOL	0.127 ***	0.281	-0.173*

Note: ***, Significance at 10%; **, significance at 5%; *, significance at 1%.

Long run estimates obtained from error correction form of exports (2) for the countries Turkey, Mexico, and Poland are shown in Table 4. According to long run estimates of the export models, the exchange rate volatility variable carries significant negative sign for Poland indicating that volatility of exchange rates lowers the export volumes. However, volatility variable has positive significant impact on real exports at ten percent significance level for Turkey. These positive and negative signs are again in line with theoretical expectations explained by “income and substitution effects” of export volumes. Second, the coefficient of the real effective exchange rate is statistically significant and negative for both Turkey and Mexico, i.e. depreciation of real effective exchange rate boosts export volumes in the long run. Finally, an increase in world income stimulates the export significantly for all three countries

as theory predicted. In other words, the results confirm the broadly accepted argument in literature that world income is the strongest determinant of a country's export.

6. Summary and Conclusions

This study investigates the export volumes of six developing countries as a function of exchange rate volatility, level of exchange rates and world income using quarterly data. Deriving the impact of exchange rate volatility on trade is first step to require to measure exchange rate volatility. There are several measures in finance literature, and I employed two of them in this study together. If exchange rate data exhibit usual features of GARCH models, such as leptokurtic distribution and negative skewness in addition to the presence of autoregressive conditional heteroskedasticity, GARCH-based proxy of exchange rate volatility may be quite reasonable. However, under the non-existence of ARCH effects, I use another measure of volatility, moving sample standard deviation. After measuring the exchange rate volatility with both GARCH and moving sample standard deviation-based proxies, ARDL-bound testing approach and error correction representation had been conducted, and long run estimations were produced.

The empirical and theoretical literature suggest that volatility of exchange rates can produce either positive or negative outcome on trade based on the magnitude of "income" or "substitution" effects. Conversely, the impact of level of exchange rate and world income on trade volumes are quite predictable. The empirical results presented in this study first reveal that there is not cointegration relationship for Brazil, Indonesia, and Chile which means that any short run deviation from the long run equilibrium level of export volumes will not correct itself over time. On the other hand, the results support the cointegration relationship for other countries, Turkey, Mexico, and Poland. Second, world income is major determinant driving export volumes in line with previous empirical studies. Third, there is a significant negative relationship between exchange rate level and export volumes for Turkey and Mexico as theory predicted. Finally, the measure of exchange rate volatility carries a significant negative coefficient in the case of Poland while there is a significant positive relationship between volatility and exports in the case of Turkey. Hence, the empirical results obtained in this study regarding with the impact of volatility of exchange rate and trade are in line with previous studies in the literature.

This study also presents some opportunities for future research. First, the results of the study can be expanded by specific trade sectors, e.g. agriculture, mining, manufacturing, energy etc. Therefore, one can evaluate the impacts of volatility of exchange rates on specific trade sectors. Second, while the exchange rate literature mostly focusses on export volumes, there are only a few studies studying exchange rate volatility on import volumes. Hence, long run import model can be estimated by the same estimation procedure conducted in this study for the case of developing countries.

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Appendix A

Table A.1: Data and Data Sources

	Period	Real Export ^a (X)	World Income ^b (W)	Real Effective Exchange Rate (RER)	Exchange Rate Volatility (VOL)
Turkey	1998:Q1 -	Nominal Export (CBRT)	Index (2010=100) of Gross Domestic Production in constant prices in OECD countries as proxy of real world income (FRED)	Real Effective Exchange Rate Index (2010=100) (FRED)	ARCH(1)
	2017:Q1	GDP Deflator (2010=100) (FRED)			
Mexico	1996:Q1 -	Nominal Export (FRED)		Real Effective Exchange Rate Index (2010=100) (FRED)	MSSD (8 Quarters)
	2017:Q1	GDP Deflator (2010=100) (FRED)			
Indonesia	2000:Q1 -	Nominal Export (FRED)		Real Effective Exchange Rate Index (2010=100) (FRED)	GARCH(1,1)
	2017:Q1	GDP Deflator (2010=100) (FRED)			
Brazil	1996:Q1 -	Nominal Export (FRED)		Real Effective Exchange Rate Index (2010=100) (FRED)	MSSD (8 Quarters)
	2017:Q1	GDP Deflator (2010=100) (FRED)			
Poland	2002:Q1 -	Nominal Export (FRED)		Real Effective Exchange Rate Index(2010=100) (FRED)	GARCH(1,2)
	2017:Q1	GDP Deflator (2010=100) (FRED)			
Chile	2000:Q1 -	Nominal Export (FRED)		Real Effective Exchange Rate Index(2010=100) (FRED)	MSSD (8 Quarters)
	2017:Q1	Consumer Price Index (2010=100) (FRED)			

Note: ^a Nominal export is deflated by GDP Deflator (2010=100) (if available) to obtain real exports for each country. CBRT and FRED are abbreviations of the Central Bank of the Republic of Turkey and Federal Reserve Economic Data, respectively.

^b I use index of total gross domestic production in constant prices for the OECD total area in line with the literature.

Table A.2: ADF Unit Root Test Results for RER

	Trend and Constant	Constant	None	Number of Lags
	Test Statistic (Critical Value)	Test Statistic (Critical Value)	Test Statistic (Critical Value)	
Turkey	-2.738 (-3.474)	-2.891 (-2.909)	-0.068 (-1.950)	1
Indonesia	-3.205 (-3.484)	-2.370 (-2.917)	-0.072 (-1.950)	2
Brazil	-2.179 (-3.468)	-1.888 (-2.904)	-0.298 (-1.950)	2
Mexico	-2.822 (-3.467)	-1.659 (-2.904)	-0.128 (-1.950)	1
Poland	-2.120 (-3.493)	-2.088 (-2.924)	-0.065 (-1.950)	3
Chile	-2.619 (-3.467)	-2.547 (-2.904)	-0.336 (-1.950)	1

Note: The null hypothesis of the ADF test is that a variable contains a unit root. Number of lags are specified by SBIC. Critical values in the parentheses are given at five percent significance level.

Table A.3: Phillips-Perron Unit Root Test Results for RER

	Trend and Constant	Constant	None	Number of Newey-West Lags
	Test Statistic (Critical Value)	Test Statistic (Critical Value)	Test Statistic (Critical Value)	
Turkey	-2.338 (-3.473)	-2.604 (-2.908)	-0.152 (-1.950)	3
Indonesia	-3.268 (-3.482)	-2.773 (-2.916)	-0.341 (-1.950)	3
Brazil	-2.086 (-3.466)	-1.868 (-2.903)	-0.298 (-1.950)	3
Mexico	-2.763 (-3.466)	-1.692 (-2.903)	-0.078 (-1.950)	3
Poland	-2.535 (-3.490)	-2.547 (-2.922)	-0.362 (-1.950)	3
Chile	-2.343 (-3.466)	-2.283 (-2.903)	-0.379 (-1.950)	3

Note: The null hypothesis of the Phillips–Perron (1988) test is that a variable contains a unit root. Test employs Newey–West (1987) standard errors to account for serial correlation. Critical values in the parentheses are given at five percent significance level. Selected Newey–West lag length is three.

