The Impact of Exchange Rate Volatility on International Trade: Evidence from the MINT Economies

Abstract

This paper examines the effect of exchange rate volatility on international trade activities for Mexico, Indonesia, Nigeria, and Turkey. We use volatility predicted from GARCH models for both nominal and real effective exchange rate data. To detect the long term relationship we use the autoregressive distributed lag (ARDL) bound testing approach; while for the short term effects, Granger causality models are employed. The results show that, in long term, there is no linkage between exchange rate volatility and international trade activities except for Turkey and even in this case the magnitude of the effect of volatility is quite small. In the short term, however, a significant causal relationship from volatility to import/export demand is detected for Indonesia and Mexico. In the case of Nigeria unidirectional causality from export demand to volatility is found, while for Turkey no causality between volatility and import/export demand is detected.

Keywords: Exchange Rates, International Trade, Volatility, Causality

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1. Introduction

Since the collapse of the Bretton Woods system, floating exchange rates' effects on international trade and the overall economy has become a significant area of investigation. In general, exchange rate uncertainty has varied effects on the economy. For instance, the level exchange rates might have a direct effect on international trade. As the international trade prices are closely related with the exchange rate fluctuations exchange rates can affect the international trade earnings and trade volumes. Exchange rate changes also influence economic policy, for example, for countries who are adopting an inflation targeting regime, central banks have to revise expected inflation target frequently, because of changes in the level and volatility in exchange rate.

Even though theory suggests that there is a negative relationship between exchange rate volatility and international trade, see for example, Arize (1997) and Doğanlar (2002), the empirical literature suggests that this theoretical argument might not always be true, see Kroner and Lastrapes (1993) and Baum and Caglayan (2010). In light of this conflicting theoretical and empirical evidence, the main aim of our research is to investigate the interaction between exchange rate volatility which is proxied by both nominal and real effective exchange rate volatility, and international trade for the four emerging market economies Mexico, Indonesia, Nigeria, and Turkey, which is also called as MINT countries (1).

The main contributions of this paper is to use the GARCH modelling procedure combined with the autoregressive distributed lag (ARDL) bounds testing approach to examine the impact of both nominal and real exchange rate uncertainty on both the long and short run trade volumes of the MINT countries. We employ the ARDL technique as it generally provides unbiased estimates of the long-run model and produces valid t-statistics even when some of the regressors are endogenous, see Odhiambo (2009). In addition, it can be used with a mixture of I(0) and I(1) data, involves a single-equation set-up, making it simple to implement and interpret and enables different variables to be assigned different lag-lengths as they enter the model which is particularly useful when studying international trade and the impact of volatility.

The paper is organized as follows. Section 2 provides a review of the literature review. Section 3 presents the data and the empirical framework of this study. Section 4 discusses the results obtained from our econometric tests and section 5 concludes.

2. Literature Review

The expected impact of exchange rate volatility on international trade activities can be positive or negative depending on the assumptions made on issues like the presence or absence of forward markets and other hedging instruments, the modeling of traders' risk preferences, the structure of production such as the prevalence of small firms and the degree of economic integration etc – see Auboin and Ruta (2013) and Oskooee and Hegerty (2007) for good recent surveys. Most theoretical studies however, support the idea that a rise in exchange rate volatility leads to a decrease in international trade volumes. According to the models if economic agents are risk averse, increased volatility in exchange rate increases uncertainty in the market and raises the cost of conducting international trade. A critical point is that it is not volatility *per se* but rather 'unanticipated volatility' Arize (1997) that is most likely to be damaging to international trade. According to Doğanlar (2002) unpredictable changes in the exchange rate between the time of the contract and delivery increase the uncertainty for exporting firm's profits. The uncertainty will be greater if there are not enough hedging instruments McKenzie (1999). When a well-developed forward market is present then the picture is very different. In a pioneering paper Ethier (1973) shows that when firms know that their revenues depend on the future exchange rate then exchange rate uncertainty will not affect the volume of trade. Other studies suggest an indirect effect of exchange rate volatility on international trade. Viane and Vries (1992) suggest that the effects of increasing exchange rate volatility on importers and exporters might be different, since they are located on different sides of the forward contract. According to this, if the trade balance and any forward risk premium are positive, exporters will lose and importers will benefit.

The initial theoretical research suggesting that exchange rate volatility is negative for international trade was based on quite important assumptions such as perfect competition, a high degree of risk aversion, the invoicing currency used, the non-existence of imported inputs and the absence of exchange rate hedging instruments. However, authors such as Broll and Eckwert (1999) show the theoretical possibility of a positive relationship between exchange rate volatility and exports. The reason for this possibility is that as exchange rate volatility increases so does the real option to export to the world market. As such higher volatility can increases the prospective gains from international trade, this applies only for firms that are able to react flexibly to changes in exchange rates and re-allocate their products accordingly. In addition, DeGrauwe (1988) emphasizes there are income and substitution effects of volatility. If firms are risk averse, a rise in exchange rate volatility increases the expected marginal utility of exports and can lead to more exports, this is the income effect. However, if firms are not risk averse enough then firms will export less because exporting is less preferable, which is the substitution effect. Consequently, depending on the relative strengths of the income and substitution effects, the net effect of the exchange rate volatility may be positive or negative.

The extensive empirical literature supports these contradictory theoretical views, Chit *et al* (2010) examine the effects of exchange rate volatility on real exports for five emerging East Asian countries and their results suggest a negative impact. When the exchange rate movements are not fully anticipated, an increase in exchange rate volatility leads risk-averse agents to reduce their international trading activities. Similarly, Arize *et al* (2008) examine eight Latin American countries and find that that there is a negative and statistically significant long-run relationship in all cases.

On the other hand, Gotur (1985) investigates the same relationship for five industrialized economies namely the USA, Germany, France, Japan, and UK and concludes that volatility has no significant effects on trade flows. Kroner and Lastrapes (1993), utilize the M-GARCH model to examine empirically the hypothesis for the same five developed countries as Gotur (1985). Their results show that exchange rate volatility has a significant effect on export flows for all countries. However, the sign of volatility coefficients are negative for the USA and the UK but positive for France, Germany and Japan. Bredin *et al* (2003) use both aggregate and sectoral export data from Ireland to the European Union, their results suggest that exchange rate volatility has no effect in the short term, but a positive and significant effect in the long term.

Hall *et al* (2010) investigate the relationship between exchange rate volatility and trade volumes for a panel of 10 emerging market economies and 11 other developing countries using quarterly data for the period 1980-2006. Their results differ between emerging markets and developing countries. Exchange rate volatility negatively affects the exports of developing countries but has no effect on exports of emerging market economies. They argue that the more open capital markets of the emerging markets may have reduced the impact of exchange rate fluctuations on exports compared with those effects in the other developing countries.

Some economists claim that exchange rate volatility, in addition to impacting the volume of trade flows also affects the variability of trade flows. Baum and Caglayan, (2010) examine the effect of exchange rate uncertainty both on the volume and variability of trade flows. They mainly focus on bilateral trade flows between 13 developed countries over the period 1980—1998. Their results show that there is no significant relationship between exchange rate uncertainty and the volume of trade. However, their results suggest that exchange rate volatility exhibits a positive impact on the volatility of international trade flows.

Finally, Haile and Pugh (2013) apply meta-regression analysis to the existing empirical literature on the impact of exchange rate volatility on international trade and find some evidence of publication bias. They show that researchers reported results are significantly influenced both by authors' modelling strategies and by the contexts of their investigations. In particular, researchers are most likely to find an adverse impact of exchange rate volatility on international trade when using low-frequency real exchange variability and focusing on trade between less developed economies which have less hedging opportunities In addition, they find that studies using nominal exchange rate volatility are less likely to report a negative impact on trade than those using real exchange rate volatility. This is because it is only over long periods that real variability diverges from its nominal value. They also report that studies employing gravity, error-correction, and long-run cointegration modelling techniques are more likely to report a negative trade impact of exchange rate volatility.

3. Data and Methodology

To investigate the effect of exchange rate volatility on international trade, as suggested by Arize *et al* (2000) export and import demand functions for four countries are estimated, adding an uncertainty variable to the specification following Tang and Nair (2002). The export and import demand functions are as follows:

$$lx_{it} = \delta_{10} + \delta_{11} lY^*_{it} + \delta_{12} lp^x_{it} + \delta_{13} V^{real}_{it} + u_{1t} \quad (1)$$

$$lx_{it} = \delta_{20} + \delta_{21} lY^*_{it} + \delta_{22} lp^x_{it} + \delta_{23} V^{nom}_{it} + u_{2t} \quad (2)$$

$$lm_{it} = \delta_{30} + \delta_{31} lY_{it} + \delta_{32} lp^m_{it} + \delta_{33} V^{real}_{it} + u_{3t} \quad (3)$$

$$lm_{it} = \delta_{40} + \delta_{41} lY_{it} + \delta_{42} lp^m_{it} + \delta_{43} V^{nom}_{it} + u_{4t} \quad (4)$$

where x and m are export and import volume, respectively, Y^* captures the rest of the world's world income conditions and Y is for domestic income, p^x and p^m are the relative export and import prices, and lastly V^{real} and V^{nom} show the real and nominal exchange rate volatility (2). The subscripts *i* and *t* are for countries and time respectively and *l* depicts that the variable is in logarithmic form.

The export and the import variables are measured by the export and the import volume indices. As a proxy for the world demand conditions we follow Chowdhury (1993) and construct the weighted average of the GDP series of each of country *i*'s ten most important trade partners in last 10 years. The relative price variables (p^x, p^m) are defined as the ratio of export and import prices of country *i* to those ten major trading partners' export and import prices. An exchange rate volatility variable is included in the model to take into account the effects of exchange rate uncertainty. Following Gür and Ertuğrul (2012) and Baum and Caglayan (2010), the variable is created by fitting a GARCH model.

The Data Set

For our empirical investigation a monthly time series dataset was used for the period from 1995M1 to 2012M12 for the four countries of interest: Mexico, Indonesia, Nigeria, and Turkey. Export (lx_t) and import (lm_t) volume data are taken from the World Bank World Development Indicators (WDI) annual database and Organization for Economic Co-operation and Development (OECD) National Accounts database. The period 1995 to 2012 is an interesting one for these economies as it covers the start of the WTO and also incorporates the period of the financial crisis that started in August 2007. The original annual data for Nigeria is taken from WDI and indexed to 2010 and converted into monthly frequency using the "quadratic-match average" frequency conversion method. Since all data were collected in current U.S. dollars, they were deflated using the consumer price index of the USA.

All GDP data (lY_t and lY_t^*) are taken from IMF's International Financial Statistics (IFS) in quarterly forms. Firstly the data were indexed to 2010 and converted into monthly frequency through the above mentioned conversion method. To calculate the world demand condition (Y^*) for Indonesia, Nigeria, and Turkey each country's ten biggest trading partners' the seasonally adjusted GDP data is taken and 10-year weighted average of GDP is calculated (3). The source of trading partners' data is the IMF Direction of Trade Statistics (DOTS).

Relative export and import prices (lp_t^x and lp_t^m) were calculated as the ratio of the export and import price of the country to the world export and import prices which is taken from IFS (4). International trade prices for export and import are available in the IFS database in monthly form for Turkey and Mexico. Foreign trade prices data for Indonesia and Nigeria were taken from the World Development Indicators (WDI). As the WDI presents data in annual form, the data is converted into monthly form.

Real and nominal effective exchange rate data which is used to predict real and nominal exchange rate volatility series were taken from Bank of International Settlements (BIS) database for Indonesia, Mexico, and Turkey while data for Nigeria was taken from IFS. The volatility series were obtained for the real and nominal effective exchange rates from estimating Generalized Autoregressive Conditional Heteroscedasticity (GARCH) models of Bollerslev (1986).

How to measure exchange rate volatility has been extensively debated in the literature and still there is no common agreement on the best proxy to show volatility. In the literature three different measures has been used to represent volatility of exchange rates. Dell'Ariccia (1999) employs the standard deviation of the first difference of the log real exchange rate. A second measure for exchange rate volatility is the moving average standard deviation of the monthly logarithm of real exchange rate, for example, Klassen (2004) and a third measure for capturing volatility, stems from ARCH/GARCH modeling. See for example, Sauer and Bohara (2001), Clark *et al* (2004) Fidrmuc and Horvarth, (2008). We use the GARCH methodology as this is the most commonly used in recent research,

4. EMPIRICAL RESULTS

Before applying the GARCH models to capture the volatility/uncertainty of exchange rates, two pre-steps are followed. First, since GARCH modelling necessitates the data used to be stationary, the stationarity of data needs to be tested. Commonly, Augmented Dickey Fuller (ADF) tests which is suggested by Dickey and Fuller (1979) are used in literature.

The data of interest is real and nominal effective exchange rates of MINT economies. All series used in logarithmic forms. The results of ADF test results are represented in Table 1 below. Here, REX and NEX are for real and nominal effective exchange rate. The prefix of L shows that the variable is in logarithmic form and suffixes stand for the countries.

Table 1. Unit Root Test Results

Variable	Level	First Difference
$LNEX_{mex}$	-4.197** (0)	
$LREX_{mex}$	-1.975(6)	-9.703**(1)
$LNEX_{ind}$	-1.910 (1)	-11.526**(0)
$LREX_{ind}$	-3.064*(1)	
$LNEX_{nig}$	-1.765 (1)	-11.596**(0)
$LREX_{nig}$	-1.742 (0)	-15.337**(0)
$LNEX_{tur}$	-4.289**(1)	
$LREX_{tur}$	-2.372 (2)	-10.384**(1)

Notes: (a) ** and * show the significance level at 1% and 5% respectively.
(b) Lag length are in parenthesis and determined by SIC

The results in Table 1 suggest that all series are stationary in first differences. However, some series seem to be level-stationary, namely, *LNEX* for Mexico and Turkey and *LREX* for India (5). Consequently, in the analysis, these three variables are used in level forms, while the others are included in first differences.

The second step is to identify the appropriate ARIMA models to be fitted to both *LNEX* and *LREX* for every country. Results from these tests are given in Tables 2 and 3 for the nominal and the real exchange rates correspondingly. It should be noted that the Akaike (AIC) and the Schwartz (SIC) information criteria were used as a model selection tool and among different models, the one which gave minimum AIC and SIC values was chosen. Based on the results, none of the models has a serial correlation problem. On the other hand, the ARCH effects test results show that, except Nigeria's real exchange rate model, all other cases exhibit ARCH effects in their residuals. This result shows that these models have heteroscedasticity problems i.e. volatility clustering in the data.

Table 2. Fitted ARIMA (p, d, q) Models for LNEX

LNEX_	MEX	IND	NIG	TUR
AR(1)	1.151 (18.005)			0.354 (4.066)
AR(2)	-0.181 (-2.894)			0.619 (7.233)
MA(1)		0.334 (5.449)	0.179 (2.420)	1.203 (12.677)
MA(2)				0.454 (7.237)
Constant	4.726 (-53.565)	-0.008 (-1.442)	-0.003 (-1.761)	4.266 (12.218)
(p, d, q)	(2, 0, 0)	(0, 1, 1)	(0, 1, 1)	(2, 0, 2)
B-G LM Test	0.405	0.738	0.386	0.436
ARCH-LM Test	0.0003	0.0000	0.0000	0.0000

Notes
(a) t values in parentheses
(b) Null Hypothesis for Breusch-Godfrey Serial Correlation test: "No serial correlation"
(c) Null Hypothesis for ARCH-LM Heteroscedasticity test: "No ARCH effect"
(d) *p* values are presented for the tests

Table 3. Fitted ARIMA (p, d, q) Models for LREX

LREX_	MEX	IND	NIG	TUR
AR(1)	-0.513 (-13.105)	0.926 (36.926)		1.003 (8.098)
AR(2)	-0.795 (-19.477)			-0.366 (-6.205)
MA(1)	0.653 (52.603)	0.294 (4.541)		-0.711 (-5.766)
MA(2)	0.988 (117.73)			
Constant	-0.001 (-0.273)	4.453 (66.607)	-0.004 (-0.519)	0.002 (1.263)
(p, d, q)	(2, 1, 2)	(1, 0, 1)	(0, 1, 0)	(2, 1, 1)
B-G LM Test	0.555	0.622	0.980	0.771
ARCH-LM Test	0.000	0.000	0.848	0.000

Notes
(a) t values in parentheses
(b) Null Hypothesis for Breusch-Godfrey Serial Correlation test: "No serial correlation"
(c) Null Hypothesis for ARCH-LM Heteroscedasticity test: "No ARCH effect"
(d) *p* values are presented for the ARCH-LM tests

Since heteroscedasticity may have an autoregressive structure, the ARCH/ GARCH methods can be employed to model the volatility in the data. We firstly fit an ARCH/ GARCH model to the data and volatility is predicted using this model. Results for ARCH/ GARCH models are presented in tables 4 and 5, optimum lag lengths are determined by AIC and SIC mentioned earlier, these models are used to predict the volatility in nominal and real exchange rate series. Since no ARCH effects were detected in the case of Nigeria's real exchange rate, ARCH/GARCH models are only estimated for the nominal exchange rate data of Nigeria.

Table 4. Fitted GARCH (p, q) Models for LNEX

LNEX_	MEX	IND	NIG	TUR
AR(1)	1.151 (18.005)			0.354 (4.066)
AR(2)	-0.181 (-2.894)			0.619 (7.233)
MA(1)		0.334 (5.449)	0.179 (2.420)	1.203 (12.677)
MA(2)				0.454 (7.237)
Constant	4.726 (-53.565)	-0.008 (-1.442)	-0.003 (-1.761)	4.266 (12.218)
resid^2 (t-1)	0.721 (6.133)	1.149 (7.731)	0.356 (3.897)	0.436 (3.817)
resid^2 (t-2)	-0.498 (-2.233)	-0.836 (-3.60)	-0.027 (-2.302)	0.132 (2.497)
GARCH (-1)	0.745 (3.261)	0.764 (5.074)		
GARCH (-2)				
Constant	0.0005 (1.070)	0.0002 (1.694)	0.0003 (10.533)	0.0004 (10.956)
ARCH-LM Test	0.0900	0.5782	0.3537	0.4291

Notes

(a) t values in parentheses

(b) Null Hypothesis for ARCH-LM Heteroscedasticity test: "No ARCH effects"

(c) *p* values for the ARCH-LM test

Table 5. Fitted GARCH (p, q) Models for LREX

LREX_	MEX	IND	TUR
AR(1)	-0.513 (-13.105)	0.926 (36.926)	1.003 (8.098)
AR(2)	-0.795 (-19.477)		-0.366 (-6.205)
MA(1)	0.653 (52.603)	0.294 (4.541)	-0.711 (-5.766)
MA(2)	0.988 (117.73)		
Constant	-0.001 (-0.273)	4.453 (66.607)	0.002 (1.263)
resid^2 (t-1)	0.597 (5.156)	1.093 (7.315)	0.244 (3.023)
resid^2 (t-2)		-0.818 (-4.059)	
GARCH (-1)		0.792 (7.420)	0.560 (4.756)
GARCH (-2)			
Constant	0.0004 (8.942)	0.0002 (1.670)	0.0002 (2.611)
ARCH-LM test	0.010	0.589	0.1942

Notes
(a) t values in parentheses
(b) Null Hypothesis for ARCH-LM Heteroscedasticity test: "No ARCH effect"
(c) *p* values for the ARCH-LM test

After obtaining the volatility/uncertainty proxies we can move to the next step which is testing for stationarity for all variables used in models (1) to (4). Similarly with our previous analysis the stationarity of all series in the model is tested through ADF unit root tests. The results of these tests are presented in Table 6.

Table 6. ADF Results for the Variables

Country	Variable	Level	First Difference
MEXICO	Export Volume	-1.486 (0)	-9.036 (1)
	Import Volume	-1.380 (0)	-6.810 (2)**
	Domestic Demand	-2.231 (10)	-3.843 (9)**
	World Demand	-1.915 (1)	-8.589 (2)*
	Export Prices	-1.555 (0)	-13.269 (0)**
	Import Prices	-2.308 (1)	-11.492 (0)**
	Volatility NEX	-10.100 (0)**	
	Volatility REX	-9.083 (0)**	
INDONESIA	Export Volume	-0.466 (1)	-19.143 (0)**
	Import Volume	-0.178 (2)	-9.652 (1)**
	Domestic Demand	0.189 (10)	-3.177 (9)**
	World Demand	0.022 (10)	3.109 (9)**
	Export Prices	-1.218 (1)	-9.876 (0)**
	Import Prices	0.065 (1)	-10.429 (0)**
	Volatility NEX	-8.759 (0)**	
	Volatility REX	-8.105 (5)**	
NIGERIA	Export Volume	-0.602 (2)	-6.856 (1)**
	Import Volume	-1.177 (3)	-2.989 (2)*
	Domestic Demand	0.554 (10)	-3.896 (9)**
	World Demand	-0.222 (10)	-3.607 (9)**
	Export Prices	-1.188 (1)	-9.429 (0)
	Import Prices	-0.686 (1)	-9.981 (0)**
	Volatility NEX	-9.429 (0)**	
	Volatility REX		
TURKEY	Export Volume	-0.573 (2)	-15.694 (1)**
	Import Volume	-0.794 (1)	-22.695 (0)**
	Domestic Demand	-0.585 (10)	-4.017 (9)**
	World Demand	-0.991 (10)	-3.028 (9)*
	Export Prices	-2.379 (0)	-16.107 (0)**
	Import Prices	-3.206 (1)*	
	Volatility NEX	-4.174 (4)**	
	Volatility REX	-3.647 (0)**	

Notes
(a) Lag length are presented in parentheses and determined by SIC
(b) Critical values for 1% and 5% are -3.472 and -2.882 accordingly
(c) * Indicates significance at 5% level, ** Indicates significance 1% level

The results reported in **Table 6** suggest that all variables, except volatility series both for *NEX* and *REX*, are nonstationary in levels, on the other hand, in first differenced forms, all variables are stationary. In other words, lx_{it}, lm_{it}, lY_{it},Y_{it}^{*}, $\mathrm{l}p_{it}^{x}$, and lp_{it}^{m} are I(1), while

V_{it}^{real} and V_{it}^{nom} are I(0). Moreover lp_{turkey}^{m} appears to be stationary in levels. This result means that it cannot used in the traditional cointegration analysis method.

There are two basic approaches to cointegration analysis, one is the Engle and Granger (1987) two-step process and the other being the Johansen (1988) maximum likelihood reduced-rank procedure. Both methods require all explanatory variables to be integrated of order one I(1). This is necessary because according to DeVitta and Abbott (2004) in the presence of a mixture of I(0) and I(1) regressors, standard statistical inference based on conventional cointegration tests is no longer valid. However, unlike the traditional methods, the ARDL bound testing technique, see Pesaran and Shin (1999) and Pesaran *et al* (2001) does not require that all the variables of interest have to be integrated of the same order.

As we have seen from the results in Table 6, the volatility series are I(0), while other series are I(1). Thus the ARDL model is the best approach for our empirical analysis. The ARDL bound testing procedure uses the *F*-statistic for the joint significance of the estimators of the lagged levels in the model to test the null hypothesis of "no cointegration". Since, the standard *F* distribution cannot be used here, Paseran *et al* (2001) provides two asymptotic critical values: the lower value assumes that all variables are I(0), and upper value assumes that all variables are I(1). If the calculated test statistic goes beyond the upper critical value, then the null hypothesis of "no cointegration" is rejected. If it falls below the lower bound the null cannot be rejected. Finally, if the statistic falls inside the respective bounds then the cointegration test is inconclusive. Once a cointegration relationship is detected the ARDL model can be applied to investigate long run and short run link between variables.

In the first step, the lag orders on the first differenced variables in equations (3) and (4) is determined from the unrestricted models by using the Schwartz Bayesian Criterion (SBC) results which are available from the authors upon request. Having obtained optimal lag order for equations (3) and (4), the next step is to employ the bounds test to investigate a long-run relationship between the variables of interest. The results of bounds F-test are presented in Tables 7 where export volume is dependent variable and Table 8 which import volume is dependent variable.

Table 7. Bound Test Results. (Dependent variable: export volume)

Country	F statistics	Lower Critical Value 5%	Upper Critical Value 5%	Volatility Measure
Mexico	8.13 (a)	4.01	5.07	Real Exchange Rate
Indonesia	8.09 (a)	4.01	5.07	Real Exchange Rate
Nigeria	--	--	--	Real Exchange Rate
Turkey	5.60 (b)	3.25	4.35	Real Exchange Rate
Mexico	8.12 (a)	4.01	5.07	Nominal Exchange Rate
Indonesia	7.92 (a)	4.01	5.07	Nominal Exchange Rate
Nigeria	2.21	3.25	4.35	Nominal Exchange Rate
Turkey	5.56(b)	3.25	4.35	Nominal Exchange Rate

Notes (a) Unrestricted intercept and trend
(b) Unrestricted intercept and no trend
(c) * Bounds test critical values are taken from Pesaran (2001)

Table 8 Bound Test Results. (Dependent variable: import volume)

Country	F statistics	Lower Critical Value 5%	Upper Critical Value 5%	Volatility Measure
Mexico	7.37 (b)	3.25	4.35	Real Exchange Rate
Indonesia	5.86 (b)	3.25	4.35	Real Exchange Rate
Nigeria	--	--	--	Real Exchange Rate
Turkey	8.84 (a)	4.01	5.07	Real Exchange Rate
Mexico	8.36 (a)	4.01	5.07	Nominal Exchange Rate
Indonesia	5.90 (b)	3.25	4.35	Nominal Exchange Rate
Nigeria	1.99	3.25	4.35	Nominal Exchange Rate
Turkey	8.79 (a)	4.01	5.07	Nominal Exchange Rate

Notes(a) Unrestricted intercept and trend
(b) Unrestricted intercept and no trend
(c) Bounds test critical values are taken from Pesaran (2001)

For Mexico, Indonesia, and Turkey all F-values are above the upper critical value which implies that there are unique cointegration vectors for all 4 models. However, the calculated F statistics for Nigeria lower than the lower bound and hence there is no cointegration relationship among variables. Another point that needs to be emphasized is that dummy variables are included in the models. For each MINT country, a dummy is created for the 2008 global financial crisis, since the effects of this crisis are reflected almost all macroeconomic variables the dummy was set a zero before 2008 month 9 and 1 thereafter. In addition, for Turkey and Mexico one more dummy variable is created because of there is a structural break in the data of these two countries. This break point is 2000M11 for Turkey and 2000M12 for Mexico. For Turkey since there is a change in slope after November, 2000, the dummy created (D_{tur}) is multiplied by a trend variable

(D_{tur} *x trend*) and included in the models. The observation that there are structural breaks in the data is also verified by the Chow break point test results for which are available upon request. The next step is to obtain the long run estimates by estimating the ARDL models. To derive the long run estimates, first the short run ARDL-ECM model should be estimated. The lag structure for ARDL-ECM is determined on the basis of SBC model selection criteria.

Table 7. ARDL-ECM Results

Country	Model 1	Model 2	Model 3	Model 4
Mexico	-0.20 (0.00)	-0.20 (0.00)	-0.18 (0.00)	-0.20 (0.00)
Indonesia	-0.23 (0.00)	-0.23 (0.00)	-0.13 (0.00)	-0.13 (0.00)
Nigeria	--	--	--	--
Turkey	-0.30 (0.00)	-0.22 (0.00)	-0.21 (0.00)	-0.19 (0.00)

Model 1: Export demand function for real exchange rate volatility
Model 2: Export demand function for nominal exchange rate volatility
Model 3: Import demand function for real exchange rate volatility
Model 4: Import demand function for nominal exchange rate volatility
Notes (a) *p* values in parenthesis
(b) Only the error correction terms are reported

The cointegration relationship between the variables is also justified by the error correction terms (ECT). The results in Table 9 show that all error correction terms are negative and statistically significant. The ECT represents the speed of recovery to long run equilibrium. For example, in the first model for Mexico, ECT -0.20 means that any deviation from the long run equilibrium is compensated in 5 months (1/0.20). The duration is 4 and 3 months for Indonesia and Turkey respectively. The ARDL-ECM was not estimated for Nigeria because there was no long run relationship among the variables.

As a last step of detection of cointegration, one can derive the long run estimates. Results for the export functions (models 1 and 2) are reported in Tables 10 and 11 respectively. Clearly world income is particularly important for all three countries exports. This is not surprising since world income is an important part of an export demand function. The income elasticity of demand for exports is found to be less than unity for the three countries.

Table 10. Long run Estimates for Export Demand Function for Real Exchange Rate Volatility (Model 1)

Countries	V_{it}^{real}	Y_{it}^{*}	$\mathrm{l}p_{it}^{x}$
Mexico[b] (1, 0, 4, 1)	1.06 (0.095)	0.20** (0.000)	0.08** (0.000)
Indonesia[c] (2, 0, 1, 0)	0.01 (0.886)	0.60** (0.000)	0.01 (0.558)
Turkey[b] (3, 0, 0, 0)	-0.09** (0.000)	0.12* (0.046)	0.17** (0.000)

Notes

(a) * significance at 5% level, ** significance at 1% level

(b) shows the inclusion of a drift term

(c) Shows the inclusion of a drift and trend term

(d) p values in parenthesis

Table 11. Long run Estimates for Export Demand Function for Nominal Exchange Rate Volatility (Model 2)

Countries	V_{it}^{nom}	Y_{it}^{*}	$\mathrm{l}p_{it}^{x}$
Mexico[b] (3, 0, 0, 0)	-0.64 (0.281)	0.20** (0.000)	0.10** (0.000)
Indonesia[c] (2, 0, 0, 0)	0.01 (0.916)	0.62** (0.000)	0.01 (0.558)
Turkey[b] (3, 0, 0, 0)	-0.05** (0.001)	0.13 (0.074)	0.19** (0.001)

Notes

(a) * significance at 5% level, ** significance at 1% level

(b) Shows the inclusion of a drift term

(c) Shows the inclusion of a drift and trend term

(d) p values are in parenthesis

The estimated coefficient of the price variable is positive and significant for Mexico and Turkey but insignificant for Indonesia. This might be because of the magnitude of the Indonesian income elasticity. Even though the export demand is income inelastic in Indonesia, it is relatively closer to unity than the other two countries and the possible effect of price may be absorbed by income elasticity in Indonesia. On the other hand, the price elasticity of export demand appears to be inelastic for both models for Turkey and Mexico.

In addition, a few words need to be said about the sign of export price variable. Intuitively, the sign of lp_{it}^{m} is expected to be negative. However, in our long run model this appears to be positive. The reason for this might be explained by causality concept, since the long run cointegration relation does not say anything about the causality between variables. At first sight it seems unexpected for lp_{it}^m to have a positive sign, if there is a causality from lx_{it} to lp_{it}^m, it is possible to expect a positive relationship between lx_{it} and lp_{it}^m. This is because, the higher the demand for exports, the higher the price of exporting goods. Hence because there is a causality from lx_{it} to lp_{it}^m it is possible that lp_{it}^m has a positive sign.

The estimated coefficients for exchange rate volatility are negative and statistically significant only for Turkey. The estimates for Mexico and Indonesia are insignificant which implies that both nominal and real exchange rate volatility do not impact on export volumes. In addition to this the magnitude of the significant coefficients for Turkey are quite small between -0.09 and -0.05 for real and nominal exchange rate volatility, respectively.

Table 8. Long run Estimates for Import Demand Function for Real Exchange Rate Volatility (Model 3)

Countries	V_{it}^{real}	lY_{it}	lp_{it}^m
Mexico[b] (1, 0, 0, 4)	-0.24 (0.678)	0.24** (0.000)	0.06 (0.744)
Indonesia[c] (2, 0, 0, 1)	-0.13 (0.452)	0.11 (0.107)	0.09* (0.020)
Turkey[b] (4, 0, 0, 2)	-0.09** (0.001)	0.15** (0.000)	-0.07** (0.004)

Notes
(a) * significance at 5% level, ** significance at 1% level
(b) shows the inclusion of a drift term
(c) Shows the inclusion of a drift term and trend term
(d) p values are in parentheses

Table 13. Long run Estimates for Import Demand Function for Nominal Exchange Rate Volatility (Model 4)

Countries	V_{it}^{nom}	lY_{it}	lp_{it}^m
Mexico[a] (1, 4, 0, 0)	-2.49** (0.002)	0.25** (0.000)	-0.01 (0.636)
Indonesia[a] (2, 0, 0, 1)	-1.07 (0.485)	1.10 (0.104)	0.86* (0.019)
Turkey[a] (4, 0, 0, 2)	-0.05** (0.005)	0.17** (0.000)	-0.08** (0.004)

Notes
(a) * significance at 5% level, ** significance at 1% level
(b) shows the inclusion of a drift term
(c) Shows the inclusion of a drift term and trend term

(d) *p* values are in parentheses

Next we look at the import demand functions for models 3 and 4 respectively which are reported in Tables 12 and Table 13. We can see that the domestic income elasticity is positive and significant in Mexico and Turkey for both models. However, interestingly, these coefficients are insignificant for Indonesia. Moreover, without exception, the results confirm that import demand is income inelastic for Mexico and Turkey.

Relative import prices are significant and positive for Indonesia but negative for Turkey and insignificant for Mexico. The estimated coefficients for relative import prices are less than unity, thus import demand is price inelastic in two of three countries. The impact of exchange rate volatility on import demand is negative and significant in Mexico and Turkey and insignificant in Indonesia. Most interestingly, the magnitude of the effect of nominal exchange rate volatility on import demand is -1.07 for Indonesia and as large as -2.47 in the case of Mexico, which implies that import demand has a highly elastic response to nominal exchange rate volatility.

Following Granger (1988) when a pair of I(1) series are co-integrated, there must be causation in at least one direction. Having established that there is a long run relationship between variables for Mexico, Indonesia, and Turkey, the next step is to detect for the causality between the variables of interest. On the other hand, since the causality concept makes no assumptions about whether the series being considered are I(0) or I(1) we can apply the test to Nigeria as well. The findings of Granger causality test is given in Tables 14-17 (one for each country). Here, for economy of space only the causality results between lx_{it}, lm_{it}, V_{it}^{real} and V_{it}^{nom} are reported.

Table 14. Granger Causality Test Results for Mexico

Model	Null Hypothesis	F statistics	Probability
Model 1	V_{it}^{real} does not Granger Cause lx_{it}	2.094	0.126
	lx_{it} does not Granger Cause V_{it}^{real}	1.882	0.155
Model 2	V_{it}^{nom} does not Granger Cause lx_{it}	1.882	0.155
	lx_{it} does not Granger Cause V_{it}^{nom}	1.225	0.296
Model 3	V_{it}^{real} does not Granger Cause lm_{it}	13.551**	0.000
	lm_{it} does not Granger Cause V_{it}^{real}	7.370**	0.000
Model 4	V_{it}^{nom} does not Granger Cause lm_{it}	12.973**	0.000
	lm_{it} does not Granger Cause V_{it}^{nom}	2.111	0.124

** Indicates 1% significance level

For Mexico a causal relationship is detected for models 3 and 4. Accordingly, in Model 3, there is a bi-directional causality is found since both null hypothesis' are rejected. Thus, one can say that, in Mexico, real exchange rate volatility and import volume are the Granger cause of each other. However, for Model 4, a unidirectional causality from nominal exchange rate volatility to import volumes is identified. The results show that there is not any casual linkage between export volumes and real or nominal exchange rate volatility.

Table 9. Granger Causality Test Results for Indonesia

Model	Null Hypothesis	F statistics	Probability
Model 1	V_{it}^{real} does not Granger Cause lx_{it}	5.925**	0.003
	lx_{it} does not Granger Cause V_{it}^{real}	1.085	0.340
Model 2	V_{it}^{nom} does not Granger Cause lx_{it}	6.526**	0.002
	lx_{it} does not Granger Cause V_{it}^{nom}	1.091	0.338
Model 3	V_{it}^{real} does not Granger Cause lm_{it}	5.458**	0.005
	lm_{it} does not Granger Cause V_{it}^{real}	1.974	0.141
Model 4	V_{it}^{nom} does not Granger Cause lm_{it}	5.860**	0.003
	lm_{it} does not Granger Cause Vol_NEX	1.934	0.147

** Indicates 1% significance level

Unlike Mexico, for Indonesia, there is not any bi-directional causal relationship for any of the four models. Yet, for all models a unidirectional casual effect has been ascertained for both real and nominal exchange rate volatility to export- import volume.

Table 16. Granger Causality Test Results for Nigeria

Model	Null Hypothesis	F statistics	Probability
Model 2	V_{it}^{nom} does not Granger Cause lx_{it}	3.519	0.032
	lx_{it} does not Granger Cause V_{it}^{nom}	12.259**	0.000
Model 4	V_{it}^{nom} does not Granger Cause lm_{it}	0.262	0.770
	lm_{it} does not Granger Cause Vol_NEX	0.121	0.887

** Indicates 1% significance level

For the case of Nigeria since volatility has been found only for the nominal exchange rate series, the Granger test is applied only for models 2 and 4. Results for model 2 show that nominal exchange rate volatility is caused by export volume. As it was mentioned before, Nigeria's export mainly depends on oil and oil prices and therefore the demand for oil is determined by international market conditions. Hence, it is sensible to expect causality from lx_{it} to V_{it}^{nom}, but not from V_{it}^{nom} to lx_{it}.

Table 10. Granger Causality Test Results for Turkey

Model	Null Hypothesis	F statistics	Probability
Model 1	V_{it}^{real} does not Granger Cause lx_{it}	0.167	0.846
	lx_{it} does not Granger Cause V_{it}^{real}	0.397	0.673
Model 2	V_{it}^{nom} does not Granger Cause lx_{it}	0.127	0.881
	lx_{it} does not Granger Cause V_{it}^{nom}	0.902	0.407
Model 3	V_{it}^{real} does not Granger Cause lm_{it}	0.015	0.985
	lm_{it} does not Granger Cause V_{it}^{real}	1.748	0.177
Model 4	V_{it}^{nom} does not Granger Cause lm_{it}	0.045	0.956
	lm_{it} does not Granger Cause Vol_NEX	2.608	0.076

Note: ** Indicates 1% significance level

Finally examining Turkey the results suggest that there is no causal relationship between either the nominal or real exchange rate volatility and import and/or export demand. However, this result does not contradict the Granger's statement which says that if there is a cointegration relationship among variables, there has to be a causal relationship in at least one direction, between any two variables in the model. This is because there are causal relationships among the other variables in each model.

In sum, the empirical results show that there is a long run relationship among variables for Mexico, Indonesia, and Turkey but not for Nigeria and this cointegration relationship is verified by significant error correction terms. In Indonesia and Nigeria both nominal and real exchange rate volatility have a short run causal effect on import demand and only in the case of Indonesia did we find a causal relationship from volatility proxies to export demand. In addition to this, no causal relationship was found between exchange rate uncertainty and import and export demand variables in Turkey. For Nigeria, we found that nominal exchange rate volatility was Granger caused by export demand. Our results for Mexico and Indonesia differ from Arize *et al* (2000) who investigate cointegration and causal relationship among export demand, export prices, world demand for exports from the Less Developed Countries (LDCs) of interest and exchange rate volatility for 13 LDCs for 1973-1996. They verified that there is a causal relationship from exchange rate volatility to export demand, and they also show that there is a cointegration relationship between variables. However, unlike our results, they show a negative and significant relationship between uncertainty and export demand in long run.

Umaru *et al* (2013) find that nominal exchange rates in Nigeria are highly volatile and Granger caused by export demand, this is consistent with the results of our study. Doğanlar (2002), investigates the effect of exchange rate volatility on real exports for 5 emerging markets including Turkey and like our model he includes export volumes, foreign economic activity (proxied by industrial production in industrial countries), relative export prices and a proxy for exchange rate uncertainty (moving standard deviation of the growth of the real exchange rate). In line with our results, he found significant and negative long run relationship between real exchange rate volatility and export demand for Turkey. His estimate for the error correction term (-0.22) for Turkey is similar to our result (-0.30). An interesting finding is that all the error correction terms reported in Table 9 are very similar. Any deviation from the long run equilibrium is corrected within 5 months for Mexico and only 4 and 3 months respectively for Indonesia and Turkey. This result show that the long run equilibrium among variables is quite stable and corrected in a relatively short time of period.

5. Conclusions

We have tested four models for the impact of exchange rate volatility on export and import demand or the MINT economies for the period 1995-2012 using monthly data. To analyze this impact, we have used both nominal and real exchange rate volatility using the GARCH model to proxy for exchange rate uncertainty. Moreover, to detect the long run relationship among variables the bound testing ARDL approach has been employed, and the Granger causality test applied to investigate the short run behavior of the variables.

We find that in the short run especially for Mexico and Indonesia volatility affects export and import demand. Hence policy makers may find some trade benefits from intervening to stabilize currency movements. However, in the long run, exchange rate volatility has no effect on export or import demand except in the case of Turkey. Even in the case of Turkey the parameter estimates show although the volatility is negatively related with the export and import demand the magnitude of this effect is quite small and should not be of great concern to policy makers.

The empirical results presented are limited to some extent. In particular since we used the real effective exchange rate the lowest frequency for the data was monthly. Secondly, the exchange rate volatility measure used only the ARCH/GARCH models. However, there are important improvements on calculating the volatility in related literature. For instance, Gür and Ertuğrul (2012) show that a switching regime GARCH (SWARCH) model can capture the volatility better in Turkey's exchange rate data than a GARCH specification. There are also other amendments in GARCH approach such as asymmetric extension of GARCH or exponential GARCH (EGARCH) or multivariate GARCH models. Future research might concentrate on which GARCH model best captures the effect of exchange rate volatility on international trade. Finally, the quality of the export and import goods also matters. In the models estimated in this study, there is no proxy employed to capture quality which over time may have a significant effect on export and import demand.

Footnotes

(1) The term of MINT is a new classification first coined by Fidelity Investments and popularized by economist Jim O'Neill (who also presented the term BRIC) in a short article in Bloomberg View in 11th of November 2013. According to O'Neill (2013), "Mexico, Indonesia, Nigeria and Turkey all have all have very favorable demographics for at least the next 20 years, and their economic prospects are interesting."

(2) Both models are estimated for nominal and real exchange rate volatility, separately.

(3) However, for Mexico, only 5 important trading partners' GDP was used since these five countries covers more than 80% of Mexico's international trade, while this rate which is calculated with ten major trading partners, for Indonesia, Nigeria, and Turkey are 72%, 58%, and 53% accordingly.

(4) The relative prices of exports and imports are calculated in dollar terms to account for the first order exchange rate effect.

(5) Note that a rejection of stationarity would imply that PPP does not hold for effective exchange rates. This could be due to several reasons, such as transaction costs, non nonlinear patterns or exchange rate overshooting.

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