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Exchange Rate Risk and Uncertainty and Trade Flows: Asymmetric Evidence from Asia

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Abstract: Very recently, the link between exchange rate volatility and trade flows has entered into a new direction in which researchers assess the possibility of asymmetric response of trade flows to a measure of exchange rate uncertainty. We add to this literature by estimating a linear and a nonlinear ARDL model to learn about the experiences of Asian countries, i.e., Pakistan, Japan, China, Korea, Singapore, Malaysia, the Philippines, and India. Like other studies in the literature, nonlinear models yielded relatively more significant results. In some cases, while the linear models showed no significant effects of exchange rate volatility on trade flows, the nonlinear models revealed significant effects. In some other cases, the opposite was true.

Keywords: exchange rate volatility; trade flows; asymmetry effects; Asia

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1. Introduction

A body of the literature in international economics includes studies that have assessed the impact of exchange rate uncertainty measured by the volatility of exchange rates on the trade flows. Early studies in the early 1970s conjectured that, since exchange rate volatility introduces uncertainty on trade flows, the effect was expected to be negative. However, theoretical developments by De Grauwe (1988) and Perée and Steinherr (1989) revealed that the effects could also be positive, depending on the degree of risk aversion by traders. Indeed, the negative and positive effects are supported by the empirical literature.

Due to the ready availability of data for developed countries, early studies used data from industrial countries to test the effects of exchange rate volatility on trade flows of these countries. The list includes Akhtar and Hilton (1984); Kenen and Rodrik (1986); Asseery and Peel (1991); Chowdhury (1993); Arize (1997); Arize and Shwiff (1998); and De Vita and Abbott (2004). Since data became available for developing countries, some researchers shifted their attention to developing countries. The list in this group includes Medhora (1990); Bahmani-Oskooee and Ltaifa (1992); Bahmani-Oskooee (1996); Doroodian (1999); Arize et al. (2000); Sauer and Bohara (2001); Hall et al. (2010); Olayungbo et al. (2011); Serenis and Tsounis (2014); Asteriou et al. (2016); and Senadza and Diaba (2017).[1]

[1] For the latest review article, see Bahmani-Oskooee and Hegerty (2007).

The above studies assumed that the response of trade flows to a measure of exchange rate volatility is symmetric. However, recently Bahmani-Oskooee and Aftab (2017) argued and demonstrated that exchange rate volatility could have asymmetric effects on trade flows. The main reason is that traders' reaction to increased volatility could be different to their reaction to decreased volatility since changes in exchange rate volatility could induce changes in expectations.

Furthermore, there is now evidence of an asymmetric response of import and export prices to exchange rate changes (Bussiere 2013). Since any volatility measure is based on the real exchange rate which includes nominal exchange rate and prices, this could be another source of asymmetric effects. Therefore, in this chapter we try to assess the asymmetric effects of exchange rate uncertainty on the trade flows of Asian countries. Indeed, as expected, we find more significant asymmetric effects of exchange rate volatility on trade flows from nonlinear models. To that end, we introduce the models and methods in Section 2, followed by empirical results in Section 3. A summary is then provided in Section 4 and data definition and sources in an Appendix A.

2. The Models and Methods

Most of the studies cited in the previous section have identified scale variables, the real exchange rate and volatility of the real exchange rate to be the three main determinants of a country's trade flows. Therefore, we follow the literature and rely upon the following export and import demand models:

$$LnX_t = \alpha + \alpha_1 LnYW_t + \alpha_2 LnREX_t + \alpha_3 LnV_t + \varepsilon_t \quad (1)$$

$$LnM_t = \beta + \beta_1 LnY_t + \beta_2 LnREX_t + \beta_3 LnV_t + \mu_t \quad (2)$$

In specifications (1) a country's exports to the world is denoted by X and is assumed to depend on the world income (YW), the real effective exchange rate (REX), and volatility of the real effective exchange rate (V). Similarly, a country's imports (M) in Equation (2) is assumed to depend on that country's own income (Y), and again the real effective exchange rate and its volatility. Since increased income leads to more imports, we expect estimates of α_1 and β_1 to be positive.[2] As the Appendix A reveals, by way of construction, since a decline in the real effective exchange rate reflects a depreciation of domestic currency, we expect an estimate of α_2 to be negative and an estimate of β_2 to be positive. This is based on the notion that a depreciation makes exports cheaper in terms of foreign currency and imports expensive in terms of domestic currency. Finally, since exchange rate volatility could have negative or positive effects on trade flows, the estimates of estimates of α_3 and β_3 could be negative or positive.

If we estimate (1) and (2) by any method, we will only obtain the long-run estimates. In order to also arrive at the short-run effects of all exogenous variables on trade flows, a common practice is to convert (1) and (2) into error-correction models. Pesaran et al. (2001) introduce a method which provides short-run and long-run estimates in one step. We follow their approach to arrive at the following error-correction models:

$$\begin{aligned} \Delta LnX_t = a_1 + \sum_{j=1}^{n1} a_{2j}\Delta LnX_{t-j} + \sum_{j=0}^{n2} a_{3j}\Delta LnYW_{t-j} + \sum_{j=0}^{n3} a_{4j}\Delta LnREX_{t-j} + \sum_{j=0}^{n4} a_{5j}\Delta LnV_{t-j} \\ +\theta_1 LnX_{t-1} + \theta_2 LnYW_{t-1} + \theta_3 LnREX_{t-1} + \theta_4 LnV_{t-1} + \varepsilon_t \end{aligned} \quad (3)$$

$$\begin{aligned} \Delta LnM_t = b_1 + \sum_{j=1}^{n5} b_{2j}\Delta LnM_{t-j} + \sum_{j=0}^{n6} b_{3j}\Delta LnY_{t-j} + \sum_{j=0}^{n7} b_{4j}\Delta LnREX_{t-j} + \sum_{j=0}^{n8} b_{5j}\Delta LnV_{t-j} \\ +\rho_1 LnM_{t-1} + \rho_2 LnY_{t-1} + \rho_3 LnREX_{t-1} + \rho_4 LnV_{t-1} + \varepsilon_t \end{aligned} \quad (4)$$

[2] Note that these estimates could also be negative if increased income is due to an increase in production of import-substitute goods (Bahmani-Oskooee 1986).

In the above specifications, the coefficients assigned to first-differenced variables reflect short-run effects and those assigned to lagged level variables reflect long-run effects. However, to match the long-run effects with long-run models (1) and (2), estimates of θ_2–θ_4 must be normalized on $-\theta_1$ in (3) and estimates of $\rho_2 - \rho_4$ must be normalized on $-\rho_1$ in (4). However, for normalized long-run estimates to be valid, we must establish cointegration. Pesaran et al. (2001) propose applying the F test to establish the joint significance of lagged level variables as a sign of cointegration. However, in this context, the F test has new critical values that they tabulate. Since the critical values account for the degree of integration of the variables, there is no need for pre-unit root testing, and indeed variables could be a combination of $I(0)$ and $I(1)$ which is another advantage of this method.

Specifications (3) and (4) are error-correction models in which trade flows are assumed to respond to changes in any of the exogenous variables in a symmetric manner. Recently, Shin et al. (2014) have modified such models so that they could also be used to assess the possibility of asymmetric effects of any of the exogenous variables on the dependent variable. Since our main purpose is to assess the asymmetric effects of exchange rate volatility, we follow Shin et al. (2014) and separate increased volatility from declines using partial sum concepts as follows

$$\begin{aligned} POS_t &= \sum_{j=1}^{t} \Delta LnV_j^+ = \sum_{j=1}^{t} \max(\Delta LnV_j, 0), \\ NEG_t &= \sum_{j=1}^{t} \Delta LnV_j^- = \sum_{j=1}^{t} \min(\Delta LnV_j, 0) \end{aligned} \tag{5}$$

where POS_t is the partial sum of positive changes in LnV and is a new time-series variable which reflects only increased volatility. Similarly, the NEG_t is the partial sum of negative changes in volatility and reflects only declines in volatility. The next step involves moving back to Equations (3) and (4) to replace LnV_t by POS_t and NEG_t. The approach results in the following new error-correction models:

$$\begin{aligned} \Delta LnX_t = {} & c_1 + \sum_{j=1}^{n1} c_{2j} \Delta LnX_{t-j} + \sum_{j=0}^{n2} c_{3j} \Delta LnYW_{t-j} + \sum_{j=0}^{n3} c_{4j} \Delta LnREX_{t-j} \\ & + \sum_{j=0}^{n4} c_{5j} \Delta POS_{t-j} + \sum_{j=0}^{n5} c_{6j} \Delta NEG_{t-j} + \lambda_1 LnX_{t-1} + \lambda_2 LnYW_{t-1} \\ & + \lambda_3 LnREX_{t-1} + \lambda_4 POS_{t-1} + \lambda_5 NEG_{t-1} + \varepsilon_t \end{aligned} \tag{6}$$

$$\begin{aligned} \Delta LnM_t = {} & d_1 + \sum_{j=1}^{n6} d_{2j} \Delta LnM_{t-j} + \sum_{j=0}^{n7} d_{3j} \Delta LnY_{t-j} + \sum_{j=0}^{n8} d_{4j} \Delta LnREX_{t-j} \\ & + \sum_{j=0}^{n9} d_{5j} \Delta POS_{t-j} + \sum_{j=0}^{n10} d_{6j} \Delta NEG_{t-j} + \pi_1 LnM_{i-1} + \pi_2 LnY_{t-1} \\ & + \pi_3 LnREX_{t-1} + \pi_4 POS_{t-1} + \pi_5 NEG_{t-1} + \varepsilon_t \end{aligned} \tag{7}$$

Such models are commonly referred to as nonlinear ARDL models and nonlinearity originates from the method of constructing partial sum variables. Shin et al. (2014) demonstrate that both the linear and nonlinear models could be estimated by the Ordinary Least Squares methods and all models are subject to the same diagnostic tests.

Once we estimate specifications (6) and (7), we can test a few asymmetric assumptions.

First, trade flows adjust to changes in exchange rate volatility at different pace, implying that the ΔPOS variable could accept a different lag order than the ΔNEG variable. In other words, if $n4 \neq n5$ if (6) and $n9 \neq n10$ in (7), "adjustment asymmetry" will be supported. Second, if at any lag order j, the coefficient estimate attached to ΔPOS_{t-j} is different to the estimate attached to ΔNEG_{t-j}, the short-run asymmetric effects of increased volatility versus decreased volatility will be supported. However, stronger evidence of short-run cumulative or impact asymmetric effects of volatility will be established if the Wald test rejects the null hypothesis $\sum \hat{c}_{5j} = \sum \hat{c}_{6j}$ in model (6) and $\sum \hat{d}_{5j} = \sum \hat{d}_{6j}$ in model (7).

Finally, if the Wald test rejects the equality of normalized coefficient estimates attached to POS and NEG variables in both models, long-run asymmetric effects of exchange rate volatility will be established.

The null hypothesis of $\frac{\hat{\lambda}_4}{-\hat{\lambda}_1} = \hat{\lambda}_5/-\hat{\lambda}_1$ in (6) and $\frac{\hat{\pi}_4}{-\hat{\pi}_1} = \hat{\pi}_5/-\hat{\pi}_1$ in (7), i.e., if normalized long-run coefficient estimates attached to the POS and NEG variables are significantly different. Again, the Wald test will be used to this end.[3]

3. Empirical Results

In this section, we estimate the linear models (3) and (4) as well as the nonlinear models (6) and (7) for Pakistan, Japan, China, Malaysia, Korea, the Philippines, Singapore, and India using quarterly data mostly over the period 1980I-2018IV. Exceptions are noted in the Appendix A. In each model, we impose a maximum of eight lags and use Akaike's Information Criterion (AIC) to select an optimum specification. Furthermore, since different estimates and diagnostic tests are subject to different critical values, we collect them in the notes to each table and use them to identify significant estimates by * and ** at the 10% and 5% levels, respectively. We begin with the estimate of the linear export demand model (3) for each country and report the results in Tables 1 and 2.

Table 1. Short-Run Coefficient Estimates of Volatility on Exports: Estimates from the linear Export Demand Model (3).

Country	Lag Number on ΔLnV								
	0	1	2	3	4	5	6	7	8
Pakistan	−1.45 (1.78) *	12.55 (1.69) *	18.20 (1.04)	16.13 (0.01)	29.07 (0.15)	5.27 (0.76)	3.47 (0.84)	5.84 (0.03)	
Japan	6.46 (1.22)	−0.86 (0.43)	−13.20 (2.24) **	−3.74 (0.34)	−5.63 (1.20)	−4.47 (1.06)			
China	−7.23 (2.51) **	4.04 (1.02)	4.26 (0.39)	4.31 (2.45) **	−5.60 (0.98)	4.50 (0.75)			
Korea	10.19 (0.95)	−13.35 (1.21)	−9.73 (1.67) *	−12.86 (1.08)					
Singapore	14.10 (1.06)	−15.58 (1.78) *	−0.60 (0.21)	−14.52 (1.93) *	−19.73 (1.82) *				
Malaysia	1.94 (0.22)	7.06 (0.56)	1.92 (0.78)	8.91 (1.38)	−11.31 (0.92)				
Philippine	21.00 (1.82) *	−3.67 (0.64)	−3.45 (1.14)						
India	3.42 (3.15) **	2.19 (1.91) *	−1.87 (0.71)	3.27 (1.79) *	−1.68 (0.80)	0.24 (0.29)			

Notes: The critical values of standard t-test are 1.64 and 1.96 at the 10% and 5% significance levels, respectively. * indicates significance at the 10% level and ** at the 5% level.

From Table 1 which reports short-run coefficient estimates attached to volatility measure we gather that it carries at least one significant coefficient in all countries except Malaysia. However, short-run effects last into the long run significant and meaningful effects only in the cases of Pakistan, Japan, Korea, Philippines, and India.[4] While the long-run effects of exchange rate volatility on Pakistan's exports are negative, in the remaining four countries they are positive.

[3] For some other applications of these methods, see Gogas and Pragidis (2015); Durmaz (2015); Baghestani and Kherfi (2015); Al-Shayeb and Hatemi-J (2016); Lima et al. (2016); Aftab et al. (2017); Arize et al. (2017); and Gregoriou (2017).

[4] By meaningful we mean cointegration is supported by either the F test or ECM_{t-1} test under which we use normalized long-run coefficient estimates and long-run model (1) and generate the error term, and we denote it by ECM. We then go back to error-correction model (3) and replace the linear combination of lagged level variables by ECM_{t-1} and estimate this new specification after imposing the same optimum lags. If ECM_{t-1} carries a significantly negative coefficient, that will support convergence of variables toward their long-run equilibrium values. Since the test with new critical values from Pesaran et al. (2001, p. 303) is used to judge significance of the estimate, this is known as the *t*-test for cointegration.

Table 2. Long-Run Coefficient Estimates and Diagnostic Statistics Associated with Linear Export Demand Models in Table 1.

Country	Long-Run Estimates				Diagnostics						
	Constant	LnYW	Ln REX	LnV	F Stat	ECM_{t-1}	LM	RESET	CUSUM	CUSUMSQ	Adj. R^2
Pakistan	13.31 (3.32) **	−1.04 (1.70) **	2.11 (2.11) **	−11.01 (2.61) **	3.89 *	−0.18 (3.33) **	1.90	1.71	S	S	0.99
Japan	−12.33 (3.11) **	0.13 (0.17)	−4.32 (2.42) **	12.21 (2.92) **	3.89 *	−0.21 (4.61) **	2.17	1.09	S	US	0.98
China	4.42 (1.31)	1.22 (2.61) **	1.11 (0.11)	−1.03 (1.51)	1.31	−0.03 (3.08) **	1.26	3.01 *	S	US	0.99
Korea	−15.71 (1.12)	1.34 (3.51) **	−4.04 (1.87) *	14.61 (1.91) *	4.11 *	−0.13 (3.23) **	1.09	1.77	S	S	0.97
Singapore	−0.61 (0.23)	−0.21 (1.00)	−1.09 (1.10)	2.19 (1.10)	3.99 *	0.02 (4.19) **	1.54	1.21	S	S	0.98
Malaysia	1.02 (1.01)	−1.12 (1.78) *	−0.81 (0.13)	0.12 (1.02)	1.18	−0.14 (2.18) **	1.21	2.01	S	S	0.99
Philippine	−15.17 (1.16)	0.011 (1.13)	−4.16 (1.81) *	11.01 (1.79) *	1.72	−0.11 (2.09) **	1.41	1.91	S	S	0.97
India	−0.102 (0.21)	1.25 (1.82) *	−1.22 (2.23) **	1.52 (2.21) **	3.82 *	−0.18 (3.21) **	1.13	1.02	S	S	0.99

a. Number inside parentheses are absolute values of the t-ratios. *, ** indicate coefficient estimates are at the 10% and 5% levels respectively. b. The upper bound critical value of the *F*-test for cointegration when there are three exogenous variables is 3.77 (4.35) at the 10% (5%) level of significance. These come from Pesaran et al. (2001, Table CI, Case III, p. 300). c. The critical value of the *t*-test for significance of ECM_{t-1} is −3.46 (−3.78) at the 10% (5%) level when $k = 3$. These come from Pesaran et al. (2001, Table CII, Case III, p. 303). d. LM is the Lagrange Multiplier statistic to test for autocorrelation. It is distributed as χ^2 with one degree of freedom. The critical value is 2.70 (3.84) at the 10% (5%) level. RESET is Ramsey's test for misspecification. It is distributed as χ^2 with one degree of freedom. The critical value is 3.84 at the 5% level and 2.70 at the 10% level.

In addition to reporting the two tests for cointegration in Table 2, we have also reported several other diagnostic tests. To make sure that the residuals are autocorrelation-free, we have reported the Lagrange Multiplier as the LM test. It is distributed at χ^2 with one degrees of freedom since we are testing for first-order autocorrelation. As can be seen, it is insignificant in all models, rendering the residuals free from correlation. Ramsey's RESET test for misspecification is also reported. This test is also distributed as χ^2 with one degree of freedom, and it is insignificant in all countries except China. Thus, all models except one are correctly specified. We have also tested for the stability of all estimated coefficients by applying CUSUM and CUSUMSQ tests to the residuals of each model. Stable estimates are indicated by "S" and unstable estimates by "US". As can be seen, almost all estimates are stable. Finally, to infer the goodness of fit, the size of adjusted R^2 is reported. Clearly, all models enjoy good fits.

Next, we estimate the error-correction model (4) associated with the linear import demand and report the estimates in Tables 3 and 4. From the short-run results in Table 3, we can gather that exchange rate volatility has significant short-run effects on the imports of all countries except the Philippines. Short-run effects last into the long-run in the cases of Pakistan, Malaysia, and India. Only in these three countries do the volatility measure carry a significant coefficient that is supported by cointegration (see Table 4). Furthermore, while exchange rate volatility has negative effects on imports of Pakistan and Malaysia, it has positive effects on imports of India. Other diagnostics are similar to those reported in Table 2 and need no repetition.

Table 3. Short-Run Coefficient Estimates of Volatility on Imports: Estimates from the Linear Import Demand Model (4).

Country	Lags on ΔLn V								
	0	1	2	3	4	5	6	7	8
Pakistan	8.21 (1.06)	18.07 (1.89) *	−12.97 (1.31)	28.31 (2.09) **					
Japan	6.35 (1.90) *	−1.10 (1.02)	−3.08 (1.01)	−7.08 (1.32)	−3.09 (1.05)				
China	−4.53 (1.00)	7.32 (2.43) **	2.27 (1.08)	1.00 (1.19)	−7.60 (1.02)	−11.12 (1.90) *			
Korea	52.21 (1.74) *	−24.92 (0.72)	29.32 (0.89)						
Singapore	39.63 (0.90)	−53.32 (1.99) **	22.82 (0.16)	−23.98 (0.98)	−54.32 (1.44)				
Malaysia	−18.65 (1.06)	56.44 (2.32) **	28.32 (1.90) *	25.03 (1.54)	46.64 (1.87) *				
Philippine	0.34 (0.21)	6.77 (1.32)	−5.68 (0.49)						
India	4.59 (1.83) *	1.56 (0.54)	3.64 (1.76) *	3.02 (1.74) *	−0.67 (0.32)	5.23 (3.32) **	1.75 (1.23)		

Notes: The critical values of standard *t*-test are 1.64 and 1.96 at the 10% and 5% significance level respectively. * indicates significance at the 10% level and ** at the 5% level.

Table 4. Long-Run Coefficient Estimates and Diagnostic Statistics Associated with the Linear Import Demand Models in Table 3.

	Constant	Ln Y	Ln REX	LnV	F Stat	ECM_{t-1}	LM	RESET	CUSUM	CUSUMSQ	Adj. R^2
Pakistan	9.11 (3.19) *	−4.12 (1.77) *	3.33 (1.69) *	−8.18 (1.78) *	3.79 *	0.39 (4.13) **	1.00	2.01	S	S	0.97
Japan	−3.65 (1.08)	−0.18 (1.95) *	−1.53 (0.33)	3.35 (1.08)	4.16 *	−0.23 (3.22) **	1.87	2.01	S	US	0.98
China	2.21 (1.14)	0.12 (1.42)	1.12 (1.15)	−2.91 (1.04)	3.88 *	0.24 (2.93) **	1.49	1.09	S	S	0.99
Korea	−5.65 (1.63)	1.16 (2.91) **	−1.41 (1.55)	2.72 (1.52)	2.01	−0.29 (2.20) **	1.08	2.08	S	S	0.97
Singapore	−3.21 (1.11)	−1.81 (1.11)	−1.17 (1.17)	4.91 (1.16)	3.89 *	−0.22 (2.24) **	1.21	1.01	S	S	0.98
Malaysia	9.71 (1.71) *	−2.15 (2.87) **	8.52 (1.78) *	−43.21 (2.82) **	1.99	−0.12 (3.16) **	1.21	1.21	S	S	0.98
Philippine	−4.94 (1.32)	−1.12 (1.61)	−2.11 (1.11)	1.31 (1.35)	2.95	−0.21 (2.31) **	1.14	1.32	S	S	0.97
India	−3.11 (1.73) *	1.31 (1.15)	−1.71 (3.09) **	1.11 (3.31) **	4.26 *	−0.23 (2.59) **	1.72	2.90*	S	US	0.99

Notes: a. Number inside parentheses are absolute values of the t-ratios. *, ** indicate coefficient estimates are at the 10% and 5% levels, respectively. b. The upper bound critical value of the *F*-test for cointegration when there are three exogenous variables is 3.77 (4.35) at the 10% (5%) level of significance. These come from Pesaran et al. (2001, Table CI, Case III, p. 300). c. The critical value of the *t*-test for significance of ECMt−1 is −3.46 (−3.78) at the 10% (5%) level when $k = 3$. These come from Pesaran et al. (2001, Table CII, Case III, p. 303). d. LM is the Lagrange Multiplier statistic to test for autocorrelation. It is distributed as χ^2 with one degree of freedom. The critical value is 2.70 (3.84) at the 10% (5%) level. e. RESET is Ramsey's test for misspecification. It is distributed as χ^2 with one degree of freedom. The critical value is 3.84 at the 5% level and 2.70 at the 10% level.

How would the results change if we consider estimates of the nonlinear models? We begin with the estimates of nonlinear export demand model (6) that are reported in Tables 5 and 6. From the short-run coefficient estimates in Table 5, we gather that either the ΔPOS or ΔNEG carry at least one significant coefficient in all countries except Korea and Malaysia. Thus, like the linear models, exchange rate volatility has short-run effects on the exports of most Asian countries. However, what the new results reveal is that the short-run effects are asymmetric because the estimate attached to ΔPOS_{t-j} at lag order j is different than the one attached to $\Delta NEG_{t-j.}$ As for short-run cumulative asymmetric effects, the Wald test reported as Wald-SR in Panel B of Table 6, rejects equality of sum of the coefficients attached to ΔPOS_{t-j} and $\Delta NEG_{t-j.}$ variables in the cases of Pakistan, Japan, China, the Philippines, and India. Short-run asymmetric effects translate into the long-run significant and meaningful effects in all countries except China and Korea. Except these two countries, in the remaining six countries either the POS or the NEG variable carry a significant coefficient (Panel A, Table 6). It appears that increased volatility only hurts exports of Pakistan, but it boosts exports of Japan, Singapore, and India. On the other hand, decreased volatility boosts exports of Pakistan and Malaysia but it hurts exports of Japan, Singapore and the Philippines. These are clear signs of long-run asymmetric effects of exchange rate volatility which are also borne out by the Wald test reported as Wald-LR in Panel B of Table 6.

Finally, we report estimates of the nonlinear import demand model (7) for each country in Tables 7 and 8. From the short-run results in Table 7, we gather that either the ΔPOS or ΔNEG carry at least one significant coefficient in all countries except Japan. The short-run effects are asymmetric because the estimate attached to ΔPOS_{t-j} at lag order j is different than the one attached to $\Delta NEG_{t-j.}$ However, stronger evidence in favor of short-run cumulative asymmetric effects are found only in the results for Pakistan, Korea, and India. In these three countries, the Wald-SR in Panel B of Table 8 is significant which rejects equality of the sum of the coefficients attached to ΔPOS_{t-j} versus $\Delta NEG_{t-j.}$ Short-run asymmetric effects translate into the long-run significant and meaningful effects in Pakistan, Korea, Singapore, Malaysia, and India, since in these countries either the POS or the NEG variable carries a significant coefficient (Panel A, Table 8). More precisely, increase exchange rate volatility hurts imports of Pakistan and Malaysia but it boosts imports of Korea, Singapore, and India. On the other hand, decreased exchange rate volatility hurts imports of Singapore and India and boosts imports of Pakistan and Malaysia. The fact that increased volatility has significant long-run effects on Korean imports but decreased volatility does not, is a clear sign of long-run asymmetric effects which is also supported by the Wald-LR test not just for Korea but also for Pakistan, China, Singapore, and India.[5]

[5] Again, other diagnostics are similar to those reported in Table 6 for nonlinear export demand models and no need to review again.

Table 5. Short-run Coefficient Estimates Attached to ΔPOS and ΔNEG in the Non-Linear Export Demand Model (6).

Country	Lags on ΔPOS								
	0	**1**	**2**	**3**	**4**	**5**	**6**	**7**	**8**
Pakistan	2.32	13.34	14.43	13.98	17.64	−8.42	−3.24	−5.16	
	(1.02)	(0.45)	(2.43) **	(3.23) **	(1.99) **	(1.13)	(1.27)	(1.60)	
Japan	6.25	−4.43	−11.54						
	(1.78) *	(0.78)	(2.56) **						
China	−23.04	1.42	−4.6						
	(2.23) **	(0.34)	(0.75)						
Korea	19.43	23.21	−2.64	5.43					
	(0.53)	(0.46)	(0.35)	(0.11)					
Singapore	55.93	−13.33	4.42	−45.43	−42.93	6.02			
	(0.56)	(2.53) **	(0.56)	(1.42)	(1.67) *	(0.38)			
Malaysia	−8.35	14.33	12.32						
	(0.43)	(1.03)	(1.41)						
Philippine	12.72	8.04	7.63						
	(1.78) *	(0.42)	(1.25)						
India	4.63	2.34	−2.73	3.42	−1.52	2.27			
	(3.34) **	(1.69) *	(1.23)	(1.21)	(0.52)	(1.83) *			
Country	**Lags on ΔNEG**								
	0	**1**	**2**	**3**	**4**	**5**	**6**	**7**	**8**
Pakistan	6.34	12.92	19.22	15.91	23.33	−6.32	−3.32		
	(0.49)	(0.09)	(2.23) **	(3.34) **	(1.92) *	(1.30)	(1.12)		
Japan	7.37								
	(1.77) *								
China	−14.53	2.33	−4.13	−10.02					
	(2.21) **	(0.27)	(0.03)	(1.83) *					
Korea	11.23	17.53	−4.03	5.03	17.03				
	(1.23)	(0.97)	(0.39)	(1.19)	(1.01)				
Singapore	15.25	−17.65	2.38	−13.42	−17.91				
	(0.89)	(2.89) **	(0.44)	(1.82) *	(1.91) *				
Malaysia	−8.02	15.09	10.46	19.31	1.21				
	(0.68)	(0.79)	(0.29)	(1.22)	(0.12)				
Philippine	13.32	10.61	5.52						
	(1.89) *	(0.39)	(1.26)						
India	4.16	4.46	−3.73	−1.51	−2.61	1.81			
	(1.86) *	(1.63) *	(0.83)	(0.53)	(1.32)	(1.15)			

Notes: The critical values of standard *t*-test are 1.64 and 1.96 at the 10% and 5% significance level respectively. * indicates significance at the 10% level and ** at the 5% level.

Table 6. Long-Run Coefficient Estimates and Diagnostics of the Non-Linear Export Demand Model (6).

Country	Panel A: Long-Run Coefficient Estimates				
	Constant	Ln Yw	Ln REX	POS	NEG
Pakistan	−7.91 (1.91) *	−0.12 (1.51)	3.91 (2.19) **	−7.98 (1.89) *	−8.75 (1.73) *
Japan	25.42 (3.33) **	0.33 (1.25)	−2.62 (2.62) **	12.32 (2.92) **	13.01 (2.82) **
China	−1.81 (1.91) *	0.11 (2.26) **	−0.71 (1.07)	−1.63 (1.24)	−1.14 (1.05)
Korea	11.71 (1.29)	−1.71 (1.26)	−1.71 (1.06)	11.1 (1.22)	2.22 (1.21)
Singapore	7.78 (1.42)	1.13 (1.20)	−7.71 (1.71) *	4.43 (2.19) *	9.92 (2.88) **
Malaysia	−63.60 (2.15) **	−1.05 (2.59) **	13.80 (1.87) *	−3.28 (1.08)	−10.72 (2.26) **
Philippine	42.94 (1.56)	1.11 (1.21)	−1.14 (1.41)	11.41 (1.54)	11.71 (2.53) **
India	2.12 (1.52)	2.92 (1.72) *	−2.48 (1.09)	1.13 (1.69) *	1.04 (1.18)

Country	Panel B: Diagnostics							
	F	ECM_{t-1}	LM	RESET	CSM(SQ)	Adj. R^2	Wald-SR	Wald-LR
Pakistan	4.90 **	−0.09 (4.13) **	1.33	2.90	S(S)	0.99	14.27 **	3.52 *
Japan	4.29 *	−0.08 (3.47) **	1.80	1.06	S(US)	0.96	2.78 *	2.81 *
China	5.93 **	0.06 (4.08) **	1.31	1.81	S(US)	0.97	2.88 *	4.09 **
Korea	1.53	−0.29 (3.14) **	1.01	2.22	S(S)	0.98	1.60	1.33
Singapore	4.32 *	−0.07 (3.78) **	1.41	1.12	S(S)	0.97	1.71	3.54 *
Malaysia	2.22	−0.03 (3.11) **	1.11	1.71	S(S)	0.98	1.13	1.72
Philippine	0.24	−0.07 (2.14) **	1.43	4.51**	S(S)	0.97	3.03 *	9.12 **
India	1.24	−0.13 (3.18) **	2.02	1.12	S(S)	0.98	4.11 **	1.61

Notes: a. Number inside parentheses are absolute values of the *t*-ratios. *, ** indicate coefficient estimates are at the 10% and 5% levels, respectively. b. The upper bound critical value of the *F*-test for cointegration when there are three exogenous variables is 3.77 (4.35) at the 10% (5%) level of significance. These come from Pesaran et al. (2001, Table CI, Case III, p. 300). c. The critical value of the t-test for significance of ECM_{t-1} is −3.66 (−3.99) at the 10% (5%) level when $k = 4$. These come from Pesaran et al. (2001, Table CII, Case III, p. 303). d. LM is the Lagrange Multiplier statistic to test for autocorrelation. It is distributed as χ^2 with one degree of freedom. e. RESET is Ramsey's test for misspecification. It is distributed as χ^2 with one degree of freedom. f. Wald tests are distributed as χ^2 with 1 degree of freedom.

Table 7. Short-run Coefficient Estimates Attached to ΔPOS and ΔNEG in the Non-linear Import Demand (7).

Country	Lag number on ΔPOS								
	0	**1**	**2**	**3**	**4**	**5**	**6**	**7**	**8**
Pakistan	10.09 (2.06) **	11.12 (2.22) **	5.19 (1.23)	17.01 (3.93) **	19.14 (1.93) *	−10.81 (1.35)	17.71 (1.91) *	10.01 (0.19)	
Japan	3.63 (0.33)	−1.12 (1.12)	−2.72 (1.02)	−2.14 (1.54)					
China	−1.32 (1.83) *	13.83 (1.93) *	1.24 (1.10)	1.33 (1.07)	−15.55 (1.98) **	−11.61 (1.78) *			
Korea	16.12 (2.04) **	−12.52 (1.41)	9.51 (1.87) *	−10.51 (1.08)	16.01 (1.51)				
Singapore	17.81 (0.51)	−16.00 (1.12)	12.21 (1.14)	−16.31 (1.11)	−10.81 (1.31)				
Malaysia	−2.33 (1.09)	19.81 (2.12) **	18.62 (1.01)	6.91 (1.73) *	15.32 (1.23)	15.22 (1.24)			
Philippine	5.85 (1.22)	4.34 (0.14)	−0.44 (0.44)	7.24 (0.14)					
India	2.72 (1.89) *	2.93 (1.81)	1.51 (0.06)	4.22 (0.38)	−1.31 (1.29)	3.33 (2.93) **			
Country	**Lag Number on ΔNEG**								
	0	**1**	**2**	**3**	**4**	**5**	**6**	**7**	**8**
Pakistan	12.32 (1.91) *	11.51 (3.12) **	11.01 (2.12) **	11.71 (3.11) **	11.91 (1.89) *	−12.93 (0.89)			
Japan	4.62 (1.29)								
China	−6.16 (1.62)	11.15 (1.71) *	6.51 (1.25)	1.25 (1.01)					
Korea	11.11 (1.72) *	−12.81 (1.08)	11.46 (1.78) *						
Singapore	18.61 (1.17)	−12.38 (1.90) *	16.71 (1.09)	−17.81 (1.13)	−16.51 (3.62) **	−14.51 (0.61)			
Malaysia	−1.46 (1.61)	9.11 (2.39) **	13.02 (1.05)	10.81 (1.85) *	17.01 (1.50)				
Philippine	8.51 (1.05)	12.52 (1.50)	−11.20 (2.12) **	17.41 (2.02) **					
India	2.22 (1.90) *	5.21 (1.78) *	2.72 (1.12)	1.89 (1.36)	−1.79 (1.61)	2.20 (1.82) *			

Notes: The critical values of standard *t*-test are 1.64 and 1.96 at the 10% and 5% significance levels, respectively. * indicates significance at the 10% level and ** at the 5% level.

Table 8. Long-Run Coefficient Estimates and Diagnostics of the Non-Linear Import Demand Model (7).

Country	Panel A: Long-Run Coefficient Estimates				
	Constant	Ln Y	Ln REX	POS	NEG
Pakistan	−4.62 (1.49)	−2.01 (1.01)	2.09 (2.23) **	−5.62 (1.90) *	−4.74 (1.87) *
Japan	8.18 (1.84) *	−1.14 (1.87) *	−1.23 (1.11)	3.23 (1.27)	3.56 (1.37)
China	−12.83 (3.21) **	−1.62 (1.32)	1.72 (1.52)	−1.53 (1.23)	−2.13 (1.23)
Korea	5.52 (1.52)	1.21 (2.82) **	−3.22 (1.23)	5.45 (1.67) *	8.32 (1.16)
Singapore	45.43 (1.95) *	−1.25 (2.19) **	−12.6 (2.03) **	34.93 (1.69) *	45.92 (2.34) **
Malaysia	−14.93 (1.87) *	−8.16 (1.88) *	12.58 (1.77) *	−12.27 (1.96) **	−12.62 (2.35) **
Philippine	7.23 (1.81) *	1.01 (2.32) **	−2.82 (1.23)	4.34 (1.18)	4.44 (1.52)
India	−2.34 (1.35)	1.71 (1.07)	−2.13 (3.09) **	4.35 (2.90) **	2.45 (3.37) **

Country	Panel B: Diagnostics							
	F	ECM_{t-1}	LM	RESET	CSM(SQ)	Adj. R^2	Wald-SR	Wald-LR
Pakistan	5.53 **	−0.23 (4.34) **	2.25	1.03	S(S)	0.99	2.92 *	3.01 *
Japan	2.03	−0.11 (2.11) **	2.12	1.39	S(US)	0.97	1.02	2.02
China	3.32 *	−0.06 (4.12) **	1.02	2.21	S(S)	0.98	1.72	3.35 *
Korea	1.21	−0.15 (3.49) **	1.21	5.32**	S(US)	0.97	3.71 *	2.92 *
Singapore	3.78 *	−0.28 (4.24) **	1.61	1.22	S(S)	0.97	2.62	4.42 **
Malaysia	2.06	−0.29 (4.32) **	1.11	1.32	S(S)	0.98	1.12	0.54
Philippine	1.72	−0.17 (3.16) **	0.11	1.07	S(S)	0.98	1.61	0.18
India	3.99 *	−0.09 (4.43) **	0.98	3.48*	S(S)	0.98	3.31 *	3.44 *

Notes: a. Number inside parentheses are absolute values of the *t*-ratios. *, ** indicate coefficient estimates are at the 10% and 5% levels respectively. b. The upper bound critical value of the *F*-test for cointegration which come from Pesaran et al. (2001, Table CI, Case III, p. 300). c. The critical value of the *t*-test for significance of ECM_{t-1} at the 10% (5%) level when $k = 4$. These come from Pesaran et al. (2001, Table CII, Case III, p. 303). d. LM is the Lagrange Multiplier statistic to test for autocorrelation. It is distributed as χ^2 with one degree of freedom. The critical value is 2.70 (3.84) at the 10% (5%) level. e. RESET is Ramsey's test for misspecification. It is distributed as χ^2 with one degree of freedom. f. Wald tests are distributed as χ^2 with 1 degree of freedom.

4. Summary and Conclusions

Ever since 1973, when the international monetary system shifted from fixed to relatively flexible exchange rates, researchers became interested in assessing the effects of exchange rate uncertainty on trade flows. Common wisdom was that more volatility will hurt international trade. However, theoretical developments revealed that more exchange rate volatility actually could boost trade, depending on the degree of risk that traders could absorb.

Recently, there is a new direction in the link between exchange rate volatility and trade flows. Since traders' reaction to increased exchange rate volatility could be somewhat different than to decreased volatility, trade flows could respond to exchange rate volatility in an asymmetric manner. Since asymmetric analysis requires estimating nonlinear models, we can also claim that research has shifted from estimating linear to estimating nonlinear models. In the chapter, we add to this new literature by considering the experiences of eight Asian countries. The list includes Pakistan, Japan, China, Korea, Singapore, Malaysia, the Philippines, and India. Following the literature to estimate the linear models, we rely upon Pesaran et al. (2001) linear ARDL approach, and to estimate the nonlinear models, we rely upon Shin et al. (2014) nonlinear ARDL approach.

Both approaches revealed that exchange rate volatility has short-run effects on trade flows of almost all countries, though the nonlinear model revealed that the short-run effects are asymmetric. However, the long-run effects were somewhat different, and more effects were found from estimating nonlinear models. More precisely, we found that increased exchange rate volatility hurts exports of Pakistan but boosts exports of Japan, Singapore, and India. On the other hand, decreased volatility boosts exports of Pakistan but hurts exports of Japan, Singapore, and India. As for the long-run effects of exchange rate volatility on imports, we found that increased volatility hurts imports of Pakistan and Malaysia and it boosts imports of Korea, Singapore, and India. On the other hand, decreased volatility boosts imports of Pakistan and Malaysia and hurts imports of Singapore and India.

The major conclusion and policy implications of our approach is that estimates are country-specific. Different results for different countries could stem from different trade rules and regulation such as tariffs, as well as different levels of governance. Therefore, in assessing the effects of exchange rate uncertainty on trade flows, we must estimate both the linear and nonlinear models to determine which approach is appropriate for which country. For example, in the results for Pakistan, both models predict that increased exchange rate volatility will hurt Pakistani exports and imports and decreased volatility will boost them. However, this is not the case in the results for Malaysia. The linear model predicted that exchange rate volatility has no long-run effects on Malaysian exports. Like most earlier studies, if we had to estimate only the linear model, we would have stopped and concluded that exchange rate uncertainty has no long-run effects on Malaysian exports to the world. However, once the nonlinear model is estimated and nonlinear adjustment of the exchange rate volatility is introduced, the results reveal that increased volatility has no effects on Malaysian exports, but decreased volatility stimulates its exports. Thus, stabilizing its exchange rate will help Malaysia to boost its exports.[6]

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Appendix A. Data Sources and Variables Definition

Quarterly data are used to carry out the empirical analysis. Depending on data availability, different periods were used for different countries as follows:

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[6] Since we have used trade flows between each country and rest of the world, the findings could suffer from aggregation bias. Future research should use bilateral trade flows between two countries to reduce the bias or even bilateral trade flows between two countries disaggregated at commodity level, e.g., Bahmani-Oskooee and Aftab (2017).

Countries	Data Period	Countries	Data Period
Pakistan	1980q1–2017q3	Korea	1980q1–2017q3
Japan	1980q1–2018q2	Philippine	1980q1–2018q2
China	1981q1–2017q3	Singapore	1980q1–2018q2
Malaysia	1991q1–2018q4	India	1980q1–2017q3

Following sources are used to collect the required data:

a. International Financial Statistics (IFS) under the IMF
b. Organization of Economic Cooperation and Development (OECD) database.

Variables

X = Total real exports of each country to the world. Nominal exports are deflated by the unit value of exports. The data come from source a.

M = Total real imports of each country from the world. Nominal imports are deflated by unit value of imports. The data come from source a.

Y = Measure of domestic income in each country. In the absence of quarterly data for any country, we followed Bergstrom (1990) and generated quarterly data. In the absence of a GDP deflator for any country, we used CPI instead. All data come from source a.

YW = World real income is proxied by the sum of real GDP of the USA and OECD countries. The data come from sources a and b.

REX = Real effective exchange rate. The data comes from source a. Missing periods are filled with the help of interpolation.

V = Volatility measure of REX. We follow Bahmani-Oskooee and Aftab (2017) and generate this measure using a Generalized Autoregressive Conditional Heteroskedasticity (GARCH 1, 1) approach. According to GARCH, our concerned variable is random, and it follows an AR (1) process as follows

$$REX = \alpha_0 + \alpha_1 REX_{t-1} + \varepsilon_t \tag{A1}$$

where ε_t is white noise with $E(\varepsilon_t) = 0$ and $\sigma^2(\varepsilon_t) = h^2$

For the purpose of forecasting the variance of REX_t, the restricted variance of ε_t should be estimated using:

$$\begin{gathered} h^2 = \beta_0 + \beta_1 \varepsilon^2{}_{t-1} + \beta_2 \varepsilon^2{}_{t-2} + \beta_3 \varepsilon^2{}_{t-3} + \beta_4 \varepsilon^2{}_{t-4} \ldots\ldots\ldots.. + \\ \beta_q \varepsilon^2{}_{t-q} + \theta_1 h^2{}_{t-1} + \theta_2 h^2{}_{t-2} + \theta_3 h^2{}_{t-3} + \theta_4 h^2{}_{t-4} \ldots\ldots\ldots.. + \theta_p h^2{}_{t-p} \end{gathered} \tag{A2}$$

To generate the forecasted values of $h^2{}_t$, we used the GARCH (p,q) model as a measure for capturing the exchange rate fluctuations. Both Equations (1) and (2) are estimated simultaneously after applying an ARCH effect. In Equation (2), GARCH order is fixed by the significance of parameters β's and θ's. In most cases, such as ours, a GARCH (1,1) specification is adequate.

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