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Impact of Exchange Rate Volatility on Trade: Empirical Evidence for the East Asian Economies

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Abstract: We investigated the impact of real exchange rate (REX) volatility on the trade performance of six East Asian economies. Several break dates were detected in the REX series over the sample period of January 1990 to December 2008. Exchange rate volatility was determined to be inherently asymmetric in all of the currency markets. We found short-run exports to be adversely affected by REX volatility in four countries, although this response was small. For imports, REX volatility was found to have a negative effect in only two of the six countries (Singapore and Indonesia). In short, there was less evidence to support the hypothesis that currency volatility has a strong negative effect on trade flows in East Asian economies. We also addressed the possibility of endogeneity in regressors and allowed for heterogeneity across countries in the model parameters.

Keywords: ARDL, EGARCH, exchange rate volatility, export, import
JEL classification: F31, C22, C51

1. Introduction

To date, the impact of exchange rate volatility on international trade remains a subject of lively debate in the literature. From a theoretical perspective, a number of models exist to show that higher exchange rate uncertainty reduces risk-adjusted expected revenue and the incentives that risk adverse traders have to engage in foreign contract. This view supports the supposition that trade performance can be impeded by exchange rate variability, and thereby, decreasing welfare. This argument is frequently used by policymakers in East Asian economies (EAEs) for a managed or fixed exchange rate regime (Arize *et al.* 2000). On the other hand, scholars such as Sercu and Vanhulle (1992) and Broll and Eckwert (1999) claim that the outcome of exchange rate volatility depends very much on the availability of a hedging mechanism. This line of literature claims that the existence of currency hedging allows economic agents to reallocate exchange rate risk, and, hence,

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can mitigate the potential adverse impact of exchange rate volatility on international trade. Meanwhile, other theoretical models predict that an increase in exchange rate risk could have either negative or positive effects on trade volume. De Grauwe (1988), for example, stressed that the dominance of income effects over substitution effects could lead to a positive relationship between trade and exchange rate volatility.[1] This counter argument can be explained as follows. If exporters are sufficiently risk averse, an increase in exchange rate variability raises the expected marginal utility of export revenue, and therefore, induces exporters to increase exports. According to De Grauwe (1988), the impacts of exchange rate uncertainty depend on the agent's degree of risk aversion. A risk adverse exporter who worries about declining revenue may export more when risks are higher as a protection from falling revenue. On the other hand, a risk-loving agent may not take a "speculative position;" thus, there would be no effects of exchange rate volatility on the agent's trade (Arize *et al.* 2008). Therefore, there is no clear-cut relationship between the two variables from a theoretical viewpoint, and the direction, as well as the magnitude of the impact of exchange rate volatility on foreign trade, is an empirical issue (Hall *et al.* 2010).

On the empirical side, a plethora of studies over the years has examined the impact of exchange rate volatility on foreign trade. Most of the studies on this relationship have focused on the major industrialised countries and the empirical evidence is no less inconclusive than the theoretical perspectives (Bahmani-Oskooee and Hegerty 2007; McKenzie, 1999). The survey results from 32 papers by McKenzie (1999) show that a total of 785 exchange rate uncertainty coefficients were estimated, of which 522 (66%) were insignificant, 191 (24%) were negative and significant, and 72 (10%) were positive and significant. Wei (1999) suggests that high exchange rate volatility discourages trade in goods at a level much greater than is typically reported for country pairs with large trade potential, but not for country pairs with relatively low trade potential. Importantly, Wei (1999) noticed that there is no empirical support for the hypothesis that the availability of hedging instruments reduces the impact of exchange rate volatility on international trade based on large groups of countries (1,000 pairs of bilateral trading partners). The author concluded by highlighting the fact that hedging is both an imperfect and costly way to avoid exchange rate risk, especially in developing economies.[2] A similar argument was later provided by Dominguez and Tesar (2001) on the subject matter. Other researchers on the issue have found that exchange rate volatility induces trade flows (Baum *et al.* 2004); they have even found no significant relationship between the two variables (e.g., Tenreyno 2007). Meanwhile, Calvo and Reinhart (2001) put forth the view that

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[1] Besides hedging, a firm also has the option to adjust its production in response to exchange rate risk. As a result, higher exchange rate risk could stimulate trade (Sercu and Uppal 2003). For a comprehensive theoretical exposition on the impact of exchange rate volatility and trade flows, the reader may refer to Bahmani-Oskooee and Hegerty (2007).

[2] Theory predicts that the negative effect of exchange rate volatility tends to disappear with the availability of hedging instruments. Wei (1999), however, rejected this hypothesis and argued that hedging instruments are often used for speculation by currency traders. Others have argued that even in the presence of a forward market, one could expect trade to be adversely affected because the transaction cost of buying cover increases the cost of trade. Additionally, trading firms generally cannot plan the magnitude or timing of all their foreign transactions.

exchange rate volatility has a larger adverse impact on trade in developing countries than in industrialised countries. Hall *et al.* (2010), however, report that the negative effects of exchange rate volatility hold for developing countries but not in the EAEs (including Korea, Singapore, Thailand and the Philippines). As Clark *et al.* (2004: 8) noted, "It is not clear whether the major changes in the world economy over the past two decades have operated to reduce or increase the extent to which international trade is adversely affected by the fluctuation in exchange rates." The fact that the literature has not reached a consensus on the impact of exchange rate volatility on trade makes the relationship an open question that warrants further investigation using more recent data. Another contribution of this paper is that the literature has overlooked structural breaks in the volatility series due to changes in the monetary policy framework in the wake of the Asian financial crisis and the economic slowdown due to the global economic recession in the early 2000s. To fill this gap, we apply a break test advocated by Bai and Perron (1998) to locate the timing of the break dates and accommodate the changing persistence in the real exchange rates in our empirical analysis.[3]

The conflicting predictions from the theoretical models and the failure of empirical studies to provide a definitive answer on the exchange rate volatility-trade volume relationship have motivated the present study. The main objective of this paper is to investigate the effect of real exchange rate volatility on international trade in East Asian countries. In these countries, the volume of trade is sizable and the variability of trade flows can significantly affect overall economic activity. In addition to focusing on the short-run impact of real exchange rate volatility, we also look at the impact of the risk in the long run. Recent papers (e.g., Bahmani-Oskooee and Harvey 2011; Bahmani-Oskooee and Hajilee 2013) based on a single equation approach have shown that the impact of the risk tends to disappear in the long run.[4] We look at this issue by using a different econometric tool—the panel data method that accounts for both heterogeneity and endogeneity of regressors.

This article contributes to the existing literature in several ways. First, we extended the analysis to include a group of six East Asian economies, covering ASEAN-5 and South Korea and a sampling period that goes beyond the 1997 Asian financial crisis and the 2000-2002 global economic slowdown. The motivation for using this particular dataset is that the currency crisis that hit most of these Asian economies has raised considerable concern about the impact of exchange rate volatility on international trade. This historical period is also characterised by episodes of shifting nominal exchange rate regimes and major changes in monetary policies in the countries under investigation. Changes in policy regimes could have affected the data-generating process of real exchange rates, and thus, shifts in conditional variance are likely to occur. An article by Lamoureux and Lastrapes (1990), among others, have argued that volatility persistence may be overstated if structural breaks are ignored in the volatility models (see also Du and Zhu 2001 and

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[3] For more discussion on the construction of the volatility series and the significance of accommodating the changing persistence in the volatility series, see Baharumshah and Soon (2014) and the articles cited therein.

[4] Bahamani-Oskooee and Hajilee (2013), who looked at 131 industries that trade between the US and Germany, found that majority of the industries react to real dollar-euro volatility in the short run. In the long run, however, only a limited number of industries were affected by exchange rate uncertainty. They concluded that the risk effect disappears in the long run.

Herwartz and Reimers 2002). For this purpose, we adopted the multiple structural-break test proposed by Bai and Perron (1998) to detect regime shifts in the exchange rate and monetary policies. Second, in terms of the methodology, we applied the ARDL approach to accommodate the problems associated with a mixture of I(1) and I(0) in time series data. Unlike past research, we relied on the Newly-West HAC standard error in the error correction model to account for the generated variable problems as highlighted by Pagan (1986). To complement the ARDL framework, we also applied Pedroni's (2001) panel cointegration method to address the issue of heterogeneity due to structural and policy differences and the potential problem of endogeneity (Tenreyno 2007).

The rest of the paper proceeds as follows. The next section provides a brief literature review. Our methodology, including the econometric model, is outlined in Section 3 and Section 4 provides the data description and presents the empirical results. Finally, Section 5 provides the concluding remarks.

2. Literature Review

The bulk of literature using aggregate trade data from industrialised countries has reported a weak or even insignificant impact of exchange rate volatility on trade. Evidence of negative impacts of exchange rate volatility on trade has only been documented in studies that use more recent data from the flexible currency regime (Doroodian 1999; Chowdhury 1993). Still, some researchers fail to provide any significant insight into the impact of exchange rate volatility on trade flows (Pattichis 2003; De Vita and Abbott 2004; Klaassen 2004; Sigh 2004). Baum *et al.* (2004) found mixed results, a negative effect in some countries and a positive for others. The evidence of a positive relationship between exchange rate volatility and trade is found in Asseery and Peel (1991) and McKenzie and Brooks (1997), among others.[5]

In recent years, the research has shifted to the EAEs, partly due to their increasing role in the global economy and the importance of exchange rate stability in the growth process. In East Asia, for instance, a great deal of attention has been paid to this topic after the onset of the 1997 Asian financial crisis. The post-1997 Asian crisis period was characterised by high volatility in REX for the majority of the EAEs. Indonesia, Malaysia, Philippines, Singapore, Thailand and South Korea are classified as upper-income developing economies with relatively more open capital markets (IMF 2007: 206-208). All six countries followed export-oriented strategies and their monetary authority is involved in active exchange rate policies. These countries depend on exports to industrialised countries, and exports are the major driving force for economic growth. The Asian crisis was actually a macroeconomic event that started in Thailand and spread quickly to several countries in the region, and eventually to the rest of the world through lower domestic demand in the crisis-affected Asian economies.[6] During the 1997 crisis, most of these countries experienced large currency depreciation as well as a collapse of the financial

[5] Studies that used sectoral data generally obtained mixed results; see McKenzie (1998) and Saito (2004) for evidence on bilateral trade flows; DeVita and Abbott (2004) and Byrne et al. (2008) for evidence on multilateral trade flows (total sectoral trade). In these studies, the effects of exchange rate volatility are found to be negative using annual data, but there may be a positive effect or no effect when using quarterly or monthly data.

[6] Developed economies benefit from economic stability as these EAEs provide opportunities for exports.

and real sectors. This landmark event has affected the growth process in these countries, where economic performance is very much connected to exchange rate stability. After the currency turmoil, Korea, Indonesia and Singapore shifted to a pure float while Thailand and the Philippines shifted to a managed float.[7]

Turning to the EAEs, which is the focus of the current study, we find that a very limited number of studies has analysed the relationship between exchange rate volatility and trade flows. A partial list of the research between 1995 and 2010 is briefly summarised in Table 1. Except for the use of the nominal exchange rate by Kim and Lee (1996), all the other studies consider the real exchange rate to examine the issue of currency volatility. Out of the eight studies listed in Table 1, four reported a negative and significant effect of REX variability on trade. For instance, research by Doroodian (1999) and Doğanlar (2002) documented evidence of a contractionary impact of real exchange rate volatility on total export flows. Similarly, Rahmatsyah *et al.* (2002) and Siregar and Rajan (2004) provide support for the negative effects of real exchange rate volatility on bilateral trade flows of Thailand and Indonesia, respectively. Evidence of real exchange rate volatility having a positive effect on trade has been stressed by Arize (1999) and Poon *et al.* (2005), while Kim and Lee (1996) found similar results for nominal exchange rate volatility. Sauer and Bohara (2001) documented mixed results from different specifications based on the panel regression model. It should be noted that two other recent studies (not reported in Table 1) by Hayakawa and Kimura (2009) and Chit *et al.* (2010), who applied the gravity model to a panel of Asian countries, also confirm the contractionary effect of exchange rate volatility on bilateral export flows.

A recent article by Hall *et al.* (2010) is not listed in the table, but their findings, based on a larger set of countries, are worth mentioning. They observed that the real exchange rate volatilities of EAEs are much lower (5.4%) than other developing countries that are not EAEs (10.9%). According to these authors, financial market deregulation in the past two decades together with a more disciplined macroeconomic policy has led to lower real exchange rate uncertainty in the EAEs compared to other developing countries. Based on panel data from 1980 to 2006, they found that EAEs (including countries under review) do not provide support for the hypothesis that real exchange rate volatility has a negative impact on exports. Meanwhile, for the developing countries, the results confirm a negative relationship between the two variables. Bahmani-Oskooee and Harvey (2011) who looks at both the short-run and long-run effect of volatility on industry trade between Malaysia and the US highlight the fact that the short-run effects could be different from the long-run effects. They highlight the fact that the short-run effects show no consistent pattern on the trade-risk relationship, but the volatility effect tends to disappear (is insignificant) in the long run in most cases.[8] Perhaps, this finding reflects the fear of the floating theory of exchange rate stabilisation. Fear of floating can be justified as currency volatility affects

[7] Hernandez and Montiel (2002) found that the crisis countries were floating more that they did prior to the crisis, the sole exception being Malaysia, which imposed capital controls and adopted a hard peg in order to move past the dramatic crisis of 1997. As for the other East Asian counties, the shift to a more flexible exchange rate regime affected the volatility of the currencies during the post crisis era. We are grateful to the referee for raising this issue to us.

[8] It worth noting their finding that aggregate exports are largely unaffected by changes in exchange rate volatility.

Table 1. Estimations from existing literature on East Asia economies

Paper	Trade model (nominal or real exchange rate), Sample and period	Coefficient for currency volatility [volatility measure]
Kim and Lee (1996)	Model: ARCH (nominal exchange rate volatility) Sample: Korea Period: 1980M1-1993M2	Real aggregate export volume 9.847**[ARCH-M]
Doroodian (1999)	Model: ARIMA (real exchange rate volatility) Sample: Korea	Aggregate export -148.77* [GARCH]
	Sample: Malaysia Period: 1973Q2-1996Q3	-497.21* [GARCH]
Sauer and Bohara (2001)	Model: fixed and random effect panel model (real exchange rate volatility) Sample: a panel of 12 Asia developing countries, which include all the countries in the present paper except Singapore Period: 1973-1993 (annual)	Real aggregate export FE: 0.907, RE: 1.147 [ARCH] FE: -1.037, RE: 0.282 [AR] FE: -1.969, RE: 0.226 [Trend]
Arize (1999)	Model: Johansen Cointegration (trade-weighted effective exchange rate volatility and real effective exchange rate volatility) Sample: Singapore Period: 1973Q2-1997Q1	Aggregate trade 0.06* [GARCH] Aggregate Export 0.08* [GARCH]
Doğanlar (2002)	Model: Engle-Granger Cointegration (real exchange rate volatility) Sample: South Korea Period: 1980Q1-1996Q4	Aggregate export -2.24 [MASD]
	Sample: Indonesia Period: 1980Q1-1996Q4	Aggregate export -1.08 [MASD]
	Sample: Malaysia Period: 1980Q1-1994Q1	Aggregate export -0.68 [MASD]
Rahmatsyah *et al.* (2002)	Model: ARDL (nominal and real exchange rate volatilities) Sample: Thailand Period: 1970Q1-1997Q2	Export to US # nil [GARCH] (-0.0004)**,(-0.0003)*[MASD] Import from US # -0.0004** [GARCH] -0.0004* [MASD]
Siregar and Rajan (2004)	Model: Johansen Cointegration (real exchange rate volatility) Sample: Indonesia Period: 1984Q1-1997Q2	Aggregate export -150.257*[GARCH] -0.003*** [MASD] Aggregate Import 4.28 [GARCH] -0.0002 [MASD]
	Sample: Indonesia Period: 1984Q1-1997Q2	Export to Japan -29.62*** [GARCH] -0.0009*** [MASD] Import from Japan -25.89*** [GARCH] -0.0004*** [MASD]
Poon *et al.* (2005)	Model: Johansen Cointegration (real effective exchange rate volatility) Sample: Indonesia, Korea, Singapore, Thailand Period: 1973Q2-2002Q2	Aggregate export Indonesia: 0.42*** [MASD] Korea: -0.23*** [MASD] Singapore: -0.19*** [MASD] Thailand: 0.09*** [MASD]

Notes: # The estimates are from EC results, there is no long-run elasticity reported in the study; (*), (**) and (***) indicate significance at 10%, 5% and 1% levels, respectively; FE = fixed effects; RE = random effects; MASD = moving average standard deviation.

a country's trade patterns and the fear that excessive volatility will discourage other countries from engaging in trade. Furthermore, the fear of floating policies performs well in more open economies with high exchange rate pass-through(see Cavoli 2009).[9]

A real exchange rate targeting based system, which is a form of fear of floating policy, appears to be well suited for countries with high pass-through and high vulnerability because the volatility of the exchange rate is much smaller than other monetary frameworks (e.g., flexible domestic inflation targeting and flexible CPI inflation targeting).

3. Methodology

Following Arize *et al.* (2000) and others, we adopted the Armington trade model to examine the trade-exchange rate volatility relationship. Accordingly, the theoretical model assumes a constant elasticity of substitution (CES) utility function derived from domestic and foreign goods and may be written as follows:

$$U = [\gamma F^{(\theta-1)/\theta} + (1-\gamma)D^{(\theta-1)/\theta}]^{\theta/(\theta-1)} \quad (1)$$

where θ is the constant elasticity of substitution between domestic and traded goods, *F* is the trade volume (export or import volumes) with country *F*, and *D* is the volume of domestic goods. For import function, the above utility function refers to the utility of the importing country, while for the export demand the utility function refers to the utility of the importing country.[10] The budget constraint for the optimisation problem is given as:

$$Y = P^F F + P^D D \quad (2)$$

where *Y* represent national income, and P^F and P^D are foreign and domestic prices, respectively. The first order necessary condition for the utility function is:

$$\frac{F}{D} = \left[\left(\frac{\gamma}{1-\gamma}\right) \frac{P^F}{P^D} \right]^{\theta} \quad (3)$$

The above equation can be express as logarithmic form given by:

$$\ln\left[\frac{F}{D}\right] = \theta \ln\left[\left(\frac{\gamma}{1-\gamma}\right)\right] + \theta \ln\left[\frac{P^F}{P^D}\right] \quad (4)$$

For empirical analysis, equation. (4) is rewritten as the following specifications:

$$\ln IM_t = \alpha + \beta_1 \ln Y_t^D + \beta_2 \ln(P_t^F - P_t^D) + \beta_3 \sigma_t^{ex} + \xi_t \quad (5)$$

$$\ln XP_t = \alpha + \beta_1 \ln Y_t^F + \beta_2 \ln(P_t^F - P_t^D) + \beta_3 \sigma_t^{ex} + e_t \quad (6)$$

where *IM* and XP_1 represent the import and export series, respectively; Y_t^D and Y_t^F are

[9] The fear of floating in EAEs (and some developed countries) is connected to a lack of credibility in their monetary policies. Exchange rate volatility tends to be damaging to trade because trade is predominantly invoiced in US dollars and hedging facilities are limited. Credit market excess is also adversely affected by currency instability. Past studies have also shown that pass-through from exchange rate swings to inflation is higher in these countries. This latter observation suggests that if price stability is a major concern of the authorities, there is a strong tendency to cap exchange rate fluctuations to narrow bands. Many of the countries in our sample have their currencies fluctuate in much the same way as they did prior to the Asian financial crises, i.e., with limited flexibility

[10] We thank an anonymous referee for pointing out this issue. For a theoretical exposition on how the exchange rate risk is incorporated in the models, see Peree and Steinherr (1989) for a detailed discussion.

the local income and foreign income, respectively; σ_t^{ex} while represent the conditional standard deviation of the exchange rate and ξ_t and e_t are the usual residual terms. Theory predicts that the elasticity (coefficients) for the income variable β_1, and the price variable β_2 are expected to be positive and negative, respectively. The elasticity for the currency volatility β_3 can be in either sign.

Based on prediction from theory, the elasticity (coefficients) for the income variable β_1, and the price variable β_2 are expected to be positive and negative, respectively. The elasticity for the real exchange rate volatility β_3 can be in either sign. The popular Johansen and Juselius (1990) cointegration test can be used to investigate the existence of a long-run cointegrating relationship if all the variables in both the equations are I(1). However, the variables in our specifications in the above are likely to be of different orders, and so we applied the bounds testing approach proposed by Pesaran *et al.* (2001), which can be applied to models irrespective of whether the regressors are I(0) or I(1), or mutually cointegrated. The bound test technique also avoids the problems of uncertainty posed by the lack of power of standard unit root tests. The bounds test for models (1) and (2) above can be examined by the following model, respectively:

$$\begin{aligned}\Delta \ln XP_t &= a^{XP} + \sum_{i=1}^{p} b_i^{XP} \Delta \ln XP_{t-i} + \sum_{i=1}^{p} c_i^{XP} \Delta \ln Y_{t-i} + \sum_{i=1}^{p} d_i^{XP} \Delta \ln(P_t^T - P_t^D)_{t-i} \\ &+ \sum_{i=1}^{p} e_i^{XP} \Delta \ln \sigma_{t-i}^{ex} + \delta_1^{XP} \ln XP_{t-1} + \delta_2^{XP} \ln Y_{t-1} + \delta_3^{XP} \ln(P_t^T - P_t^D)_{t-1} + \delta_4^{XP} \ln \sigma_{t-1}^{ex} + \mu_t^{XP}\end{aligned} \quad (7)$$

$$\begin{aligned}\Delta \ln IM_t &= a^{IM} + \sum_{i=1}^{p} b_i^{IM} \Delta \ln IM_{t-i} + \sum_{i=1}^{p} c_i^{IM} \Delta \ln Y_{t-i} + \sum_{i=1}^{p} d_i^{IM} \Delta \ln(P_t^T - P_t^D)_{t-i} \\ &+ \sum_{i=1}^{p} e_i^{IM} \Delta \ln \sigma_{t-i}^{ex} + \delta_1^{IM} \ln IM_{t-1} + \delta_2^{IM} \ln Y_{t-1} + \delta_3^{IM} \ln(P_t^T - P_t^D)_{t-1} + \delta_4^{IM} \ln \sigma_{t-1}^{ex} + \mu_t^{IM}\end{aligned} \quad (8)$$

where Δ is the first difference operator, u_t is the residual term and all the variables of the models are defined earlier. The null hypothesis of non-existence of the long-run relationship is tested using the F-test: $H_0 : \delta_1 = \delta_2 = \delta_3 = \delta_4 = 0$ against the alternative hypothesis H_0 at least one of the δ is not zero.

The critical value bounds of the F-statistics for different numbers of regressors (k) are tabulated in Pesaran and Pesaran (2003). Two sets of critical values are provided. The upper bound assumes that all the variables in the ARDL model are I(1) while the lower bound assumes all variables to be I(0). Cointegration is confirmed irrespective of whether the variables are I(1) or I(0) if the computed F-statistic falls outside the upper bound, and rejected if outside the lower bound. Nevertheless, if the F-statistic falls within the critical value band, a unit root test of stationarity is needed to authenticate the order of integration of respective variables. If these series are indeed found to be cointegrated, an unrestricted error correction version of the corresponding ARDL models can be estimated to trace the short-run dynamic of the model. The unrestricted error correction model is given by a general form of the error correction model (ECM) as follows:

$$\Delta \ln XP_t = -\phi(1,\hat{p})EC_{t-1}^{XP} + \sum_{i=1}^{k} \beta_{i0} \Delta x_{it} + \delta' \Delta w_t - \sum_{j=1}^{\hat{p}-1} \phi_j^* \Delta \ln XP_{t-j} - \sum_{i=1}^{k} \sum_{j=1}^{\hat{q}_i-1} \beta_{ij}^* \Delta x_{i,t-j} + \mu_t \quad (9)$$

$$\Delta \ln IM_t = -\phi(1,\hat{p})EC_{t-1}^{IM} + \sum_{i=1}^{k} \beta_{i0} \Delta x_{it} + \delta' \Delta w_t - \sum_{j=1}^{\hat{p}-1} \phi_j^* \Delta \ln IM_{t-j} - \sum_{i=1}^{k} \sum_{j=1}^{\hat{q}_i-1} \beta_{ij}^* \Delta x_{i,t-j} + \mu_t \quad (10)$$

with $$EC_t^{XP} = \ln XP_t - \sum_{i=1}^{k} \hat{\theta}_i x_{it} - \hat{\psi}' w_t \tag{11}$$

$$EC_t^{IM} = \ln IM_t - \sum_{i=1}^{k} \hat{\theta}_i x_{it} - \hat{\psi}' w_t \tag{12}$$

where x are the i number of regressors, i.e. ln Y_t, ln$(P_t^{I} - P_t^{D})$ and lnσ_t^{ex}, while w captures the drift and trend components, and $\hat{p}$ and $\hat{q}$ the lag length. Accordingly, the long-run coefficient for all the explanatory variables can be obtained by normalising the dynamic short-run coefficients of each explanatory variable to one minus the short-run coefficients of the dependent variable.

Previous studies show that there are no qualitative differences in using the nominal or real exchange rate (see McKenzie and Brooks (1997)), but in this paper, our focus is on the real exchange rate volatility. We define the real exchange rate as below:

$$REX_i = ln[NEX_{i/j}(P_i/P_j)]$$

where NEX is the nominal exchange rate and P_i and P_j are the price level at country i and j, respectively.

To begin, we employed the endogenous multiple structural break tests advocated by Bai and Perron (1998) (BP, hereafter) to detect possible breakpoints in the data-generating process of the real exchange rate series. The break tests revealed that two breaks were fairly common in the REX series for the countries under investigation.[11] Two breaks were detected for Korea (March 1997 and December 1999), Indonesia (May 1997 and December 2001), Philippines (April 1997 and December 1999) and Singapore (April 1997 and December 1999). As expected, the first break date was closely connected to the 1997 Asian financial crisis, while the second break coincided with the adoption of inflation targeting and the changing monetary framework adopted in most of the crisis-affected countries. For Malaysia, the break test results revealed three structural breaks (January 1997, February 2000 and August 2005) in the REX series, with the third break date occurring around the abolishment of the hard peg and the lifting of capital controls. There is high heterogeneity between series regarding the timing of the break dates detected by BP procedure, although major financial crisis seem to coincide with most of them. This could be connected to the various interventions in the foreign exchange market (capital controls) or shifting in the monetary anchor (inflation targeting). We should take into account these break points in the empirical analysis that follows. These break dates are used to construct dummy variables for estimating exchange rate volatility. To conserve space, the results of the BP tests are not reported here but are available upon request.

We applied the Nelson's (1990) Exponential Generalised Autoregressive Conditional Heteroskedasticity (EGARCH hereafter) to generate the series of conditional exchange rate volatility. The model has several added advantages over the conventional GARCH model in generating the REX volatility. The asymmetric EGARCH model is able to simultaneously accommodate asymmetric volatility and the leverage effect of bad news in the currency market. A more important merit of this model is its ability to capture large shocks of any

[11] Recently, researchers have suggested that many macroeconomic series might contain two or more breaks.

sign in financial series, which is particularly critical given the currency crash period in our sample. The model is specified in the following equations:

$$REX_t = \mu + \delta REX_{t-1} + \varepsilon_t + \varphi\varepsilon_{t-1} \text{ where } \varepsilon_t|\Omega_{t-1} \sim \text{GED}(r) \tag{13}$$

$$\log\sigma_t^2 = \omega + \sum_{i=1}^{p}\left(\alpha_i\left|\frac{\varepsilon_{t-i}}{\sqrt{\sigma_{t-i}}}\right| + \gamma_i\frac{\varepsilon_{t-i}}{\sqrt{\sigma_{t-i}}}\right) + \sum_{j=1}^{q}\beta_j\log\sigma_{t-j}^2 + \sum_{k=1}^{b}\eta_k Break_k \tag{14}$$

Equations (13) and (14) are the conditional mean and variance of the logarithm of monthly REX, respectively. To remove linear dependency from the mean equation, the conditional mean is assumed to follow an Autoregressive Moving Average (ARMA) process of ARMA(1,1); i.e., with one lag dependent variable and one lag error term. The log transformation in the variance equation (10) ruled out negative variance, so no restriction is needed in the variance equation to ensure a positive volatility process as in the conventional GARCH model. In the conditional mean equation (13), μ is the intercept term; δ and φ represents the magnitudes of the autoregressive and the moving average terms, respectively; and ε_t is the idiosyncratic news. In the conditional variance equation (14), σ_{t-1}^2 represents the lagged conditional variance of ε_t while α, β and γ are the parameters of ARCH, GARCH and leverage parameters, respectively. The response of REX (conditional volatility) to good and bad news are asymmetric if $\gamma \neq 0$, but symmetry if $\gamma = 0$. The presence of the leverage effect can be tested by the hypothesis of $\gamma < 0$. The order of EGARCH (p=1,q=1) is sufficient to capture the dynamics of the financial time series data (Bollerslev *et al.* 1992: 10). It is widely acknowledged that the EGARCH model presupposes the Gaussian assumption. Given the overwhelming evidence of non-normality in macroeconomic data, it is unlikely that the conventional GARCH model is able to adequately proxy exchange rate uncertainty. We follow Nelson (1991) to use a more general distribution—the generalised error distribution (GED)—which is more flexible and includes several other distributions (e.g., the Laplace, normal, student-t and double exponential).

The dummy variable $Break_k$ with the parameters η_k represent the structural breaks effect in the volatility process, especially the one during the Asian financial crisis. By entering these dummy variables into the volatility model, we can capture the effects of changing exchange and monetary regimes as mentioned earlier. We found that when regime shifts in the condition variances are considered, the persistence of REX volatility is much lower. In other words, accommodating for the breaks actually weakens the persistence in volatility to some degree. This result is also consistent with the finding in earlier studies that have argued that exchange rates are determined not only by fundamental factors but also, to a major extent, by the subjective perception of market participants (see Morales-Zumaquero- and Sosvilla-Rivero 2014).

4. Data and Empirical Results

Our analyses on the trade-risk relationship are based on monthly frequency data spanning from January 1990 to December 2008 for six East Asian countries, namely Indonesia, Malaysia, the Philippines, Singapore, Thailand and South Korea. Early studies on these countries are based on data of a different frequency spanning from the 1970s and 1980s up to the period prior to the 1997 crisis. These sampling periods rule out the effects of the shifting exchange rate regime and monetary policy adopted by these countries in the post-crisis period. We should note that all six countries had already embarked on financial market liberalisation. We choose not to go beyond 2008 as the recent subprime crisis that started in 2009 has yet to be concluded.

All of our data were downloaded from DataStream and the trade series (import and export) were extracted from the IMF Direction of Trade Statistics (DOTS). The time series data are nominal exchange rates quoted in USD, import and export values, industrial production to proxy for income, and consumer price index (CPI). CPI was used to proxy domestic and foreign prices, which was used to compute the real exchange rate series REX. We conducted a seasonal adjustment on both the export and import series with a multiplicative moving average method. It should be noted that all variables used in the analysis are expressed in real terms.

The EGARCH estimates for REX volatility are not reported here to conserve space. Basically only the Malaysian ringgit and Singapore dollar have a leverage effect, but they are statistically insignificant.[12] This outcome is largely due to the control of the structural break points in the variance equation.[13] Nearly all the estimates for the structural break parameter were positive and highly significant, implying that the volatilities were significantly exaggerated during the endogenous detected structural distortions. The diagnostic checking implies no major deficiency in our exchange rate fitted models.

Table 2 summarises the descriptive statistics for the series used in the empirical analysis. The statistics reported are the sample size, mean standard deviation, Jarque-Bera normality test and the Kwiatkowski–Phillips–Schmidt–Shin (KPSS 1992) unit root tests as well. The signs for all the mean values of the variables are as expected. The mean of the exchange rates is positive, indicating that, on average, the Asian currencies are depreciating over the sample period being examined. The conditional standard deviation is estimated from model (14) as a proxy for REX volatility with a value smaller than zero; as a result, it is a negative value when expressed in logarithmic form. As expected, almost all the variables (including exchange rates) are non-normally distributed. This is possibly due to the presence of large extreme values, high excess kurtosis, and time varying behaviour inherent in the exchange rate series. As discussed earlier, the application of the GED method can account for these non-normalities. Focusing on the KPSS test, we notice that the stationary null is rejected in all but two variables (volatility series) when the test contains an intercept only. However, the results favour a mixture of I(0) and I(1) when the test statistic contains both an intercept and a deterministic trend. Notice that for South Korea, both the dependent variables are found to be stationary. These results should be interpreted with some caution because of the well-known power and size problems. For robust checks, we apply the Perron and Ng (2001) unit root test, which is well known for its power and size, and the result is shown in Table 3. Our findings based on the Ng-Perron test are supportive of the unit root null and are in line with earlier studies. Given that none of these variables follow the I(2) process, we proceed with Pesaran's ARDL modelling.[14]

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[12] Tse and Tsui (1997) reported leverage effects exist for the Malaysian ringgit, but not for the Singapore dollar.

[13] In our preliminary estimates that use a shorter sample period (1990-2005) without the structural break dummies, the asymmetry parameters were all significant, although with a positive sign, indicating no leverage effects in these East Asian currencies.

[14] The authors are grateful to a referee of this journal for bringing this issue to our attention.

Table 2. Descriptive statistics and unit root tests

Variable	Mean	Std dev	Normality	KPSS in level	KPSS first difference			
				Intercept	Trend and intercept	Intercept	Trend and intercept	Stationary level
Export Series (XP)								
Korea	7.8238	0.3920	18.9302*	1.4077*	0.0956	0.0849	0.0453	I(0)
Indonesia	6.6096	0.4501	23.7772*	0.5033*	0.1912*	0.1875	0.1875*	I(1)
Malaysia	7.3590	0.5610	30.4602*	1.5500*	0.3555*	0.3732	0.0910	I(1)
Philippines	6.5435	0.4182	26.6634*	1.4867*	0.4457*	0.2711	0.0621	I(1)
Singapore	7.1802	0.2328	19.5107*	1.6620*	0.2386*	0.1122	0.0505	I(1)
Thailand	7.0341	0.4314	12.3681*	1.2279*	0.1710*	0.5000*	0.5000*	I(1)
Import Series (IM)								
Korea	7.5322	0.3050	11.7492*	1.8917*	0.1117	0.0962	0.0920	I(0)
Indonesia	5.4630	0.3548	7.9224*	1.7139*	0.3496*	0.1279	0.0678	I(1)
Malaysia	6.5458	0.4104	39.4975*	1.7339*	0.3609*	0.6134	0.0674	I(1)
Philippines	6.1568	0.4429	42.3859*	1.2487*	0.4848*	0.5680	0.0829	I(1)
Singapore	7.1880	0.3587	4.0916	0.6615*	0.3501*	0.2998	0.0602	I(1)
Thailand	6.1470	0.3493	3.0532	1.8503*	0.2988*	0.3500	0.0677	I(1)
Industrial Production (IP)								
Korea	4.1430	0.4243	14.9625*	1.9843*	0.0625	0.3461	0.1177	I(0)
Indonesia	4.4530	0.1829	29.0948*	1.4087*	0.2351*	0.1969	0.0922	I(1)
Malaysia	4.1897	0.3873	15.5109*	1.9350*	0.3426*	0.3064	0.0547	I(1)
Philippines	4.4436	0.4600	16.5289*	1.9370*	0.3943*	0.4907*	0.0980	I(1)
Singapore	4.0514	0.3549	8.0352*	1.9290*	0.1004	0.4882*	0.4485*	I(0)
Thailand	4.5833	0.3835	13.3342*	1.9029*	0.3109*	0.0872	0.0785	I(1)
US	4.5117	0.1674	24.6072*	1.8557*	0.4006*	0.3388	0.1102	I(1)
Price Differential (PD)								
Korea	0.0711	0.0792	36.0760*	1.7352*	0.4552*	0.5309*	0.0750	I(1)
Indonesia	0.5344	0.5217	25.6813*	1.9548*	0.1974*	0.0822	0.0822	I(1)
Malaysia	-0.0097	0.0248	1.5319	0.5995*	0.4174*	0.1050	0.1034	I(1)
Philippines	0.1947	0.1988	18.5173*	1.9083*	0.4223*	0.4940*	0.1129	I(1)
Singapore	-0.1004	0.0780	21.1932*	1.9409*	0.3116*	0.2122	0.1871*	I(1)
Thailand	0.0347	0.0625	25.1952*	1.4330*	0.3928*	0.3618	0.1056	I(1)
Exchange Rate Volatility (EXV)								
Korea	-4.1421	0.7237	9.8675*	0.7534*	0.2508	0.1321	0.0732	I(0)
Indonesia	-3.5969	0.9929	16.8265*	0.7794*	0.3161	0.0833	0.0809	I(0)
Malaysia	-4.3754	0.9709	15.7633*	0.2372	0.1562	0.0651	0.0641	I(0)
Philippines	-3.7901	0.4727	116.2867*	0.1790	0.1543	0.0230	0.0233	I(0)
Singapore	-4.1468	0.2937	22.2184*	0.5446*	0.1788	0.0546	0.0475	I(0)
Thailand	-4.1451	0.7234	11.3279*	0.7545*	0.2457	0.1044	0.0613	I(0)

Note: (*) indicates rejection of null at the 5% significance level. The Kwiatkowski-Phillips-Schmidt-Shin (KPSS) unit root test is based on the assumption the series is stationary or trend stationary under the null hypothesis. The stationary level is determined based on the lowest order of rejection.

Table 3. Perron and Ng (2001) unit-root test

	Constant					Constant and trend					
	MZa	MZt	MSB	MPT	k	MZa	MZt	MSB	MPT	k	
Imports											
Korea	0.16	0.13	0.81	40.53	3	-3.26	-1.15	0.35	25.39	13	I(1)
Indonesia	-4.58	-1.40	0.31	5.58	5	-10.24	-2.26	0.22	8.90	5	I(1)
Malaysia	-1.69	-0.88	0.52	13.78	4	-7.84	-1.73	0.22	12.28	3	I(1)
Philippines	-0.30	-0.24	0.81	36.37	2	-5.91	-1.49	0.25	15.17	2	I(1)
Singapore	-0.22	-0.21	0.94	47.47	3	-2.47	-0.83	0.34	26.93	3	I(1)
Thailand	-0.75	-0.49	0.66	23.83	4	-6.56	-1.67	0.25	13.97	4	I(1)
Exports											
Korea	-1.04	-0.67	0.64	21.26	6	-3.54	-1.16	0.33	23.08	6	I(1)
Indonesia	0.83	0.86	1.04	72.10	14	-3.55	-1.30	0.36	25.05	13	I(1)
Malaysia	0.20	0.18	0.90	48.76	10	-6.61	-1.74	0.26	13.85	10	I(1)
Philippines	0.17	0.19	1.10	68.98	6	-1.22	-0.42	0.34	31.73	5	I(1)
Singapore	-0.21	-0.22	1.02	55.09	6	-0.17	-0.09	0.53	64.50	4	I(1)
Thailand	0.90	1.79	1.98	249.07	2	-2.18	-0.82	0.37	31.17	2	I(1)

Note: The lag (k) selection is based on modified AIC with maximum lag 14.

We first need to establish cointegration among the variables in the model to infer the long-run effects. For this purpose, we imposed a maximum of 12 lags on each first difference variables in both the import and export equations, and then used the Akaike Information Criterion (AIC) to select the optimum lag length; see Bahmani-Oskooee and Goswami (2003) for details. Overall, the computed F-statistics for the estimated equations are greater than the upper bound of the critical values at the 5 per cent significance level or better. Clearly, the joint significance of the lagged level variables implies a long-run relationship between exports (imports) and its determinants. The long-run coefficients of REX, the variable of our focus, are all insignificant in most cases at standard significance levels in both of the estimated equations.[15] For brevity of presentation, we chose not to report these results, but are available upon request.

A potential problem in the regressions reported so far is that the exchange rate volatility variable may be endogenous. Governments may choose to stabilise the exchange rate with important trading partners. Following Arize *et al.* (2008) and others, we take a further step in our analysis by specifically addressing the endogeneity between exchange rate volatility and trade by using the fully modified ordinary least squares (FMOLS) estimator of Phillips and Hansen (1990). The results are presented in Table 4. Several interesting features emerge from this table. First, we find five countries (ASEAN-5: Singapore, Malaysia, Thailand and Philippines) where exchange rate volatility has a significant negative effect on export growth while the effect for Korea is insignificant (zero). Second, we find that in four countries (Singapore, Thailand, Indonesia and Philippines) the trade effect is either negative or positive for import growth while in the other two, it is statistically insignificant even at the 10 per cent level. There seems to be heterogeneity with regard to the impact of exchange rate risk on trade, reflecting the potential policy and structural differences in the

[15] More recent studies that consider both the short- and long-run relations have reported similar results.

individual countries. It is unclear if the size of trade has any bearing on the negative effect of volatility. It is also unclear whether hedging instruments may benefit these countries. Finally, the large income elasticities (recorded in some countries) appear to endorse the result found in past studies (see Arize *et al.* 2008). This confirms the importance of foreign economic activities on economic progress in the region.

Turning to the ECM models, the single equation diagnostic statistics in Tables 5 and 6 indicate that the estimated models perform reasonably well. First, the LM statistics indicate that autocorrelation is not a problem in any of the estimated equations. Second, there is no evidence of heteroscedaticity as indicated by Engle's LM ARCH test. Third, there is some evidence of non-normality in two equations (Singapore's exports and Thailand's imports). Finally, all of the estimated equations, with the notable exception of the equation for Singapore's imports, pass Ramsey's RESET test, indicating that the ECM is adequately specified. Note that the figures reported in parentheses in Tables 4 and 5 are the Newey-West HAC standard error to address Pagan's (1986) critique regarding the problems of the so-called generated variable. Readers should take note when comparing our results with those reported in previous literature.

We can now make several observations regarding the results summarised in Tables 5 and 6. First, the majority of short-run coefficient estimates are significant at the 10 per cent significance level or better. Second, the results show that changes in foreign income in the export equation have a positive impact on real exports in all six countries, as predicted by theory. Third, the cumulative sum of the coefficients of the exchange rate volatility is negative in all countries for the export equations; the exception is Indonesia. We conducted the Wald test of joint significance of the lagged REX volatilities, but the results are not reported here to conserve space. The cumulative sum was found to be significantly different from zero. The negative sign in the export equation implies that an

Table 4. FMOLS long run estimate

	Panel A: Export (to US)			Panel B: Import (from US)		
	IP(US)	PD	EXV	IP(US)	PD	EXV
Countries specific coefficients						
Indonesia	3.78a	0.37a	-0.06a	1.59a	0.15c	-0.09a
	[8.55]	[2.72]	[-2.97]	[7.16]	[1.80]	[-2.90]
Korea	-0.15	-1.98c	0.06	0.59a	-0.81	0.04
	[-0.33]	[-1.92]	[1.31]	[4.87]	[-1.11]	[0.86]
Malaysia	2.37a	3.11a	-0.06a	0.74a	2.20b	0.03
	[26.76]	[5.05]	[-4.04]	[11.13]	[2.09]	[1.34]
Philippines	0.65	-2.19a	-0.11b	0.78a	-0.03	0.09c
	[1.27]	[-5.11]	[-2.39]	[2.89]	[-0.04]	[1.65]
Singapore	4.64a	-5.82a	-0.10c	1.94a	-4.13a	0.22c
	[20.96]	[-12.66]	[1.65]	[7.19]	[-3.40]	[1.76]
Thailand	2.37a	-0.71	-0.09b	0.32b	-3.39a	-0.15b
	[7.31]	[-0.68]	[-2.17]	[2.54]	[-3.32]	[-2.10]
Panel group	2.28a	-1.20a	-0.06b	0.99a	-1.00c	0.02
	[26.34]	[-5.15]	[-4.87]	[14.60]	[-1.62]	[0.25]

Notes: (a), (b) and (c) indicate significance at the 1%, 5% and 10% levels, respectively. The values in [ ] denote the *t*-statistic. IP(US)=US industry production, PD=relative price, and EXV=exchange rate volatility with breaks by ARMA(1,1)-EGARCH(1,1).

Table 5. Error correction model for export

	Indonesia		Korea		Malaysia		Philippines		Singapore		Thailand	
Intercept	0.0144a	(0.0045)	-0.0011	(0.0061)	0.0013	(0.0045)	0.0064	(0.0047)	0.0010	(0.0031)	0.0054	(0.0033)
Δln[XP(-1)]	-0.4092a	(0.0690)	-0.6301a	(0.0672)	-0.2969a	(0.0548)	-0.2936a	(0.0698)	-0.3165a	(0.0587)	-0.4720a	(0.0524)
Δln[XP(-2)]	-0.1958a	(0.0690)	-0.3864a	(0.0655)			-0.1939a	(0.0527)			-0.2747a	(0.0517)
Δln[XP(-3)]					0.2862	(0.0588)			0.2561a	(0.0679)		
Δln[XP(-4)]							0.1851a	(0.0618)	0.1599a	(0.0595)		
Δln[XP(-5)]			-0.1345a	(0.0467)								
Δln[IP)]	0.2905	(0.2234)	0.5291a	(0.2535)	0.2411	(0.2004)	0.7059a	(0.2562)	0.6305a	(0.1738)	0.4005b	(0.1568)
Δln[IP(-1)]			0.9156a	(0.3168)			0.9360a	(0.2640)	0.2847	(0.1756)	0.7177a	(0.2095)
Δln[IP(-2)]			0.9451b	(0.3899)			0.4895b	(0.2027)			0.6000a	(0.2282)
Δln[IP(-3)]			0.5378c	(0.3051)							0.3205c	(0.1684)
Δln[PD]	0.3126	(0.2498)	0.2812	(0.8349)			-0.6489	(0.4644)			-1.0859	(0.4168)
Δln[PD(-2)]									1.6359b	(0.6383)		
Δln[PD(-3)]					-2.0437a	(0.7516)						
Δln[EXV]					-0.0171	(0.0124)	0.0013	(0.0110)	0.0202	(0.0314)		
Δln[EXV(-1)]	0.0204a	(0.0098)					-0.0126	(0.0134)	-0.0261	(0.0340)		
Δln[EXV(-2)]					-0.0238c	(0.0129)	-0.0219c	(0.0125)				
Δln[EXV(-3)]							-0.0068	(0.0125)				
Δln[EXV(-4)]			-0.0470c	(0.0241)			-0.0105	(0.0105)	-0.0341	(0.0317)	-0.0659a	(0.0164)
Δln[EXV(-5)]					0.0159	(0.0119)	-0.0062	(0.0124)			-0.0276b	(0.0132)
ECT(-1)	-0.1450a	(0.0349)	-0.0455c	(0.0284)	-0.2251a	(0.0460)	-0.0537c	(0.0281)	-0.1359a	(0.0354)	-0.0608b	(0.0253)
Diagnostic Checking												
LM(4)	1.0022	[0.4074]	1.7030	[0.1101]	0.4545	[0.7691]	0.4840	[0.7475]	1.5451	[0.1539]	1.1254	[0.3456]
ARCH (7)	1.1441	[0.3368]	0.7469	[0.6324]	0.8620	[0.5375]	0.1728	[0.9905]	1.376	[0.2169]	1.4883	[0.1729]
Norm (2)	3.8620	[0.1450]	8.0078b	[0.0182]	2.5693	[0.2767]	1.1188	[0.5715]	0.9157	[0.6326]	0.7496	[0.6874]
HET	1.4465	[0.1473]	0.7944	[0.7186]	0.9156	[0.5522]	1.3337	[0.1338]	1.1566	[0.2961]	0.9750	[0.4959]
RESET	0.3362	[0.7149]	0.7017	[0.4969]	0.8330	[0.4362]	3.9824b	[0.0201]	0.0297	[0.9707]	0.4324	[0.6495]

Notes: (a), (b), and (c) denote significance at the 1%, 5% and 10% levels, respectively. The values in ( ) and [ ] refer to Newey-West HAC standard error and *p*-value, respectively. The equation summary statistics include the LM test for fourth-order correlation, the Engle's (ARCH) test for heteroskedasticity, the Jarque-Bera (Norm) test for normality, White's heteroscedasticity test (HET) and Ramsey's RESET test.

Table 6. Error correction model for import model

	Indonesia		Korea		Malaysia		Philippine		Singapore		Thailand	
Intercept	0.0096	(0.0131)	0.0113	(0.0080)	0.0067	(0.0065)	0.0154c	(0.0092)	0.0061	(0.0055)	-0.0025	(0.0118)
Δln[IM(-1)]	-0.3039a	(0.0727)	-0.4373a	(0.0740)	-0.5132a	(0.0813)	-0.4539a	(0.0600)	-0.4610a	(0.0689)	-0.5062a	(0.0840)
Δ ln[IM(-2)]	-0.1503a	(0.0569)	-0.2529a	(0.0721)	-0.3388a	(0.0896)	-0.1512b	(0.0604)	-0.1999a	(0.0476)	-0.3041a	(0.0905)
Δ ln[IM(-3)]			-0.1903a	(0.0638)	-0.2158b	(0.0912)			-0.0983c	(0.0585)	-0.2259a	(0.0737)
Δ ln[IM(-4)]					0.0413	(0.0610)					-0.1569b	(0.0689)
Δ ln[IP)]	0.0054	(0.1283)	0.1438	(0.1263)	0.2924c	(0.1474)	-0.1384	(0.1101)	0.0566	(0.0490)	0.1784	(0.1743)
Δ ln[IP(-1)]					0.2810	(0.1743)	-0.1685	(0.1206)				
Δ ln[IP(-3)]			0.0820	(0.1130)								
Δ ln[PD]	2.3048b	(1.004)	-0.0912c	(1.6410)					0.8252	(1.2510)		
Δ ln[PD(-1)]			1.1890	(1.3530)								
Δ ln[PD(-3)]					1.5834	(1.0310)						
Δ ln[PD(-4)]					2.8090b	(1.1890)						
Δ ln[PD(-5)]			-0.3159	(1.262)			3.1086b	(1.4010)			5.1376b	(2.0930)
Δ ln[EXV]					0.0110	(0.0166)						
Δ ln[EXV(-1)]	0.0464b	(0.0231)					-0.0412	(0.0263)			0.1329a	(0.0498)
Δ ln[EXV(-2)]					-0.0215	(0.0199)	0.0383	(0.0281)				
Δ ln[EXV(-3)]					-0.0224	(0.0187)						
Δ ln[EXV(-4)]	-0.0600b	(0.0244)	-0.0406	(0.0362)								
Δ ln[EXV(-5)]			-0.0377	(0.0410)	-0.0169	(0.0138)	0.0394c	(0.0213)	-0.0980b	(0.0455)		
ECT(-1)	-0.2021a	(0.0684)	-0.1715a	(0.0505)	-0.1173a	(0.0399)	-0.1348a	(0.0428)	-0.0598b	(0.0265)	-0.1239b	(0.0537)
Diagnostic Checking												
LM(4)	2.9766b	(0.0204)	0.4281	(0.7883)	0.8690	(0.5321)	0.3792	(0.8234)	1.1781	(0.3216)	2.0367c	(0.0522)
ARCH (7)	0.6108	(0.6553)	0.7913	(0.5954)	1.1455	(0.3360)	1.2029	(0.3025)	1.0531	(0.3953)	0.8355	(0.5589)
Norm (2)	2.0962	(0.3506)	1.0844	(0.5815)	0.6549	(0.7208)	3.9647	(0.1377)	2.4594	(0.2924)	1.7803	(0.4106)
HET	1.4435c	(0.0612)	1.3684	(0.1300)	1.5291c	(0.0543)	0.4557	(0.9728)	1.1284	(0.3328)	0.6431	(0.8460)
RESET	0.3243	(0.7234)	0.3822	(0.6828)	0.2494	(0.7795)	0.0348	(0.9658)	2.1315	(0.1213)	1.5483	(0.2151)

Notes: (a), (b), and (c) denote significance at the 1%, 5% and 10% levels, respectively. The values in ( ) and [ ] refer to Newey-West HAC standard error and *p*-value, respectively. The equation summary statistics include the LM test for fourth-order correlation, the Engle's (ARCH) test for heteroskedasticity, the Jarque-Bera (Norm) test for normality, White's heteroscedasticity test (HET) and Ramsey's RESET test.

increase in exchange rate uncertainty reduces exports. Focusing on the import model (Table 6), our results revealed that REX volatilities were negative in Indonesia and Singapore but were somewhat positive in Thailand and the Philippines. The outcome rejects the notion that exchange rate uncertainty adversely affects imports in the Asian countries, with the notable exception of Singapore and Indonesia.[16] Again, this finding reveals that there are no "one size fit all" answers with regard to the impact of exchange rate uncertainty on trade because the fundamental result is that these markets react differently to exchange rate risk.

Finally, the coefficient estimate of the error correction term (ECT_{t-1}) is negative and statistically significant in all countries. The coefficient estimates (in absolute value) range from 0.0455 to 0.2251 and 0.0598 to 0.2021 for exports and imports, respectively. A significant negative ECT_{t-1} reinforced the cointegrating relationship based on the F-test discussed earlier. These findings confirm the long-run relation among the variables in each equation. Focusing on the export equation, this means that when exports exceed their long-run relationship with foreign income, relative prices, and exchange rate volatility, they adjust at a rate of 5 to 22 per cent per month. Likewise, we observed that imports adjust to any shocks in all the countries and the speed of adjustment varies from one country to another. We should mention that most of the test statistics from the CUSUM and CUSUM squares tests do not reject the null hypothesis of parameter stability at the 5 per cent level.[17]

Given our special interest in the impact of the REX volatility on trade flows in the countries examined, it is important to discuss the role of this variable in some detail. The magnitude and the sign on the short-run impact of the variable vary across countries. The hypothesis that exchange rate volatility has an adverse effect on trade flows is rejected by the data in the cases of Malaysia and Thailand, given the impact does not differ significantly from zero in both the import and export equations. A striking feature of our results is that the estimated impact of real exchange rate uncertainty is small in comparison with most prior studies (see Table 1). Bahmani-Oskooee and Harvey (2011) found that the short-run effects do not last in the long run in most cases, but we found that this is not so for the Asian countries, meaning that it is significant only in the short-run relationship.[18] Finally, we added a dummy variable in the regression as an exogenous variable to capture the impact of the 1997 crisis and the move towards a more flexible exchange rate regime. However, the coefficient of the dummy variable turned out to be statistically insignificant. This result (not reported) seems to suggest that an adverse effect of macroeconomic uncertainty during the crisis is adequately captured by the exchange rate volatility variable.

To summarise, the estimates of the short-run dynamics of the ECM indicate that

[16] For comparison with other countries, Arize *et al.* (2008), for example, found that exchange rate volatility has a consistent negative impact for the Latin America countries ranging from -0.40 to -0.001.

[17] To conserve space, the plots of both tests are not reported here but are available from the authors upon request. For comparison, we also applied the panel cointegration tests using four statistics suggested by Pedroni (2001). In all, the computed statistics suggest the null of no cointegration is easily rejected for the panel of countries under review. These rresults are available from the authors upon request.

[18] Readers may refer to Bahmani-Oskooee and Harvey (2011) and the papers cited therein on the implication of a variable that is significant in the short run but not in the long run.

exchange rate volatility has significant short-run negative effects on export (import) demand in only four (two) countries.[19] In countries where we found volatility to have a negative (or positive) effect on trade, the effect appears to be small. One reason could be the growing share of international transactions undertaken by multinational firms (MNCs) in the counties under review. With their presence in the domestic economy, the exchange rate risk impact may have a declining impact on world trade because fluctuations in different exchange rates may have an offsetting effect on their profitability.

5. Conclusions

We found the volatility model to fit the data well and confirm that exchange rate volatilities are affected by major economic events, including the 1997 Asian currency crisis of late 1990s. The post-crisis era is characterised by higher exchange rate volatility in all but one country, Malaysia, which pegged the ringgit to the US dollar (September 1998-July 2005). For the other Asian countries, the finding is consistent with the view that the exchange rate regime is more flexible as they increasingly integrate with the global financial system. Neglecting structural breaks may lead to misspecification of the conditional variance, casting serious doubt on the reliability of the findings from the trade-exchange rate volatility relationship.

Neither theory nor empirical studies provide us with a firm answer on the effect of exchange rate volatility on trade and consequently on the overall health of the economy. We found statistical impact of real exchange rate volatility to be significant in some, but not all of the countries examined. Our results highlight that real exchange rate volatility has a significant negative effect on exports in the short-run for South Korea, Malaysia and Philippines. Meanwhile, we found volatility to have a negative effect on the volume of imports in the short run in two countries (Indonesia and Singapore) but it turned out to be positive in Thailand and the Philippines. It is reasonable to conclude that some evidence of a dampening effect of real exchange rate volatility on trade flows is found in some, but not all the countries. In fact, our results reveal that the East Asian countries appear to be unevenly affected by real exchange rate volatility. The economic impact, however, is small in relation to the relative magnitude of the other core variables—relative price and income. Based on the panel data method, Hall *et al.* (2010) found that the impact of real exchange rate volatility on exports in EAEs differs from other developing countries. In contrast, our findings highlight that its impact may vary from one country to another. Much the same can be concluded for exports. Hall *et al. (*2010) show that exchange rate volatility has no significant (negative) impact on exports in the East Asian economies. On the other hand, our study seems to endorse the view that export performance is affected by exchange rate volatility.

The theoretical argument put forward in the literature suggests that exchange rate

[19] We also estimated the long-run equations using the ARDL approach. The income elasticity is all-significant at the 1% level and is in the elastic range. For REX volatility, the low *t*-values (Indonesia, sole exception) make it difficult to interpret. Thus, these findings as well as the evidence presented in some recent studies appear to suggest that a negative relationship is not supported by the data in the long run. To conserve space, these results are not reported here but are available from the authors upon request.

volatility has an indeterminate impact on exports and imports. Our empirical investigation supports this view for a group of Asian countries that are highly dependent on exports to boost economic growth. There are no "one size fit all" answers with regard to the impact of exchange rate uncertainty on trade flows. In some of these countries, we find that exchange rate volatility does not adversely affect their exports or imports. Their capital markets are able to adapt to movements in exchange rate regimes. This also implies that macroeconomic policies aimed at stabilising exchange rates are unlikely to increase the volume of trade in some countries.

Finally, our analysis is based on bilateral aggregate trade flows data. The use of aggregate data assumes that income, price and exchange rate elasticities are equal across the sectors of the economy. A further extension would consider a selection of important industries in these Asian countries. Modelling trade flows at sectoral levels could broaden our understanding of the relationship between exchange rate uncertainty and trade performance across sectors of the economy in recent decades where increasing integration with the world economy through trade has taken place. We may need to consider a link between exchange rate volatility and foreign direct investment (FDI) to understand whether a move towards an East Asian currency market could increase international trade.

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