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# EXCHANGE RATE VOLATILITY AND TRADE FLOWS OF EAST ASIAN EMES[ℜ]

*HOOY CHEE WOOI, AHMAD ZUBAIDI BAHARUMSHAH*

*ABSTRACT*

This paper investigates the impact of real exchange rate (REX) volatility on the trade performance of six East Asia emerging market economies. Several break dates were detected in the data generating process (DGP) of REX and volatility series over the sample period 1990:1 to 2008:12. We applied the bounds testing approach developed by Pesaran et al. (2001) and the Pedroni (2001) panel cointegration tests and find that trade flows have a long-run relationship with income, price differential and exchange rate volatility. Overall, the long-run relationship reveals that trade flows are income elastic, but not price sensitive. In contrast to the existing literature, we found that trade is responsive to exchange rate risk, albeit this response is small. For two countries, we find the exchange rate fluctuation have no effect or even a positive effect on international trade.

Keywords: Exchange rate volatility, Export, Import, EGARCH, ARDL

## INTRODUCTION

To date, the impact of exchange rate volatility (or variability) on international trade flows remains lively debated 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 of risk adverse traders 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 some policymakers in the emerging market economies (EMEs, hereafter) for a managed or fixed exchange rate regime. On the other hand, scholars such as Sercu and Vanhulle (1992), Broll and Eckwert (1999), and others claim that the outcome of exchange rate volatility depends very much on the availability of a hedging mechanism. This line of literature argues that the existence of currency hedging allows agents to reallocate exchange rate risk, and hence, can mitigate the potential adverse impact of exchange rate volatility on international trade. Meanwhile, other theoretical models predict that an increase in risk could be expected to have either negative or positive effects on trade volume. De Grauwe (1988), for example, stressed that the dominance of income effects over substitution effects leads to a positive relationship between trade and exchange rate uncertainty. 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" and thus there would be no effects of exchange rate volatility on the agent's trade (see also Arzie 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 (Arize et al., 2000; Hall et al., 2010).

On the empirical side, over the years a large number of studies have examined the impact of exchange rate volatility on foreign trade. Most of the studies on this relationship have focused on the major industrialized countries and the empirical evidence is no less inconclusive than the theoretical perspectives (Bahmani-Oskooee and Hegerty, 2007; McKenzie, 1999). For example, the empirical article by Wei (1999)

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suggests that high exchange rate volatility discourages goods trade at a level that much greater than is typically reported for country pairs with large trade potential, but not for country pairs with relatively low trade potential. Additionally, the author pointed out 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 (i.e., based on 1000 pairs of bilateral trading partners).[1] A similar argument was made by Dominguez and Tesar (2001). Other researchers on the issue have found that exchange rate volatility induces trade increases (see, for example, Baum et al., 2004), or have even found no significant relationship between the two variables. Meanwhile, other scholars such as Calvo and Reinhart (2001) put forth the view that exchange rate volatility has a larger adverse impact on trade in developing countries than in industrialized countries. Hall et al. (2010), however, report that the negative effects of exchange rate volatility hold for the developing countries but not in the EMEs.[2] As Clark et al. (2004, p.8) stated "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 existing literature has not reached a consensus on the impact of exchange rate volatility on trade makes the relationship an open empirical question that warrants further investigation.

Research in recent years has shifted to the EMEs, partly due to their increasing role in the global economy and the importance of stability of exchange rate in the growth process. In East Asia, a great deal of attention has been given to this topic after the onset of the 1997 Asian financial crisis. The post 1997 Asian crisis period was characterized by high volatility in real exchange rates (REX) of East Asian EMEs. Indonesia, Malaysia, Philippines, Singapore, Thailand and South Korea are classified to be upper-income developing economies with relatively more open capital markets (IMF, 2007, pp. 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 the industrialized countries, and exports are the driving force for economic growth. The 1997 Asian crisis was actually a macroeconomic event began 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.[3] During the 1997 crisis, most of these countries experienced large currency depreciation as well as a collapse of the financial and real sectors. This landmark event has affected the growth process in these countries, where growth performances are 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 manage float.[4]

The conflicting predictions from the theoretical models and the failure of the empirical studies to provide a definitive answer have motivated the present study. The main objective is to investigate the effect of exchange rate volatility on international trade in the emerging East Asian EMEs. This article contributes to the existing literature in several ways. First, we extend the analysis to include a group of six East Asian EMEs, 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 periods. The motivation for using this particular data set is that the currency crisis that hit most of these Asian economies in 1997 has raised concern about the impact of exchange rate volatility on international trade. This historical period is also characterized by episodes of shifting nominal exchange rate regimes and major changes in monetary policies. Changes in policy regimes could have affected the DGP of exchange rates, and thus, shifts in conditional variance may be likely. The articles by Lamoureux and Lastrapes (1990), among others, have argued that volatility persistence may be overstated if structural breaks are ignored in the volatility models

[1] According to the hedging hypothesis, the negative effect of exchange rate volatility tends to disappear (reduced) with the availability of hedging instruments. Wei (1999) rejected the 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 increase the cost of trade. Additionally, trading firms cannot always plan the magnitude or timing of all their foreign transactions.

[2] The panel of 10 emerging market economies consists of Argentina, Brazil, Hungry, Israel, Korea, the Philippines, Singapore, South Africa, Thailand and Turkey.

[3] The developed economies benefit from the economic stability as these emerging market economies provide opportunities for exports.

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

(see also Du and Zhu, 2001 and Herwartz and Reimers, 2002). Ignoring the possibility of structural breaks may affect the risk measurement and erroneously show the presence of a negative relationship between trade flow and exchange rate variability. For this purpose, we adopt a multiple structural-break test proposed by Bai and Perron (1998, 2003) to detect regime shifts in the exchange rate and monetary policies. Thus, our study expands earlier research by considering the breaks due to changing policy regimes in the empirical analysis. Second, in terms of research methodology, we applied the ARDL approach and relied on the Newey-West HAC standard error in error correction model to account for the generated variable problems as highlighted by Pagan (1986). We also conducted the Pedroni (2001) panel cointegration method to show the relationship between trade flows and its determinant. The long-run parameters of the model are estimated using the group-means fully modified OLS (FMOLS) of Pedroni (2001), which are designed to handle serial correlation and allows for heterogeneity across the individual countries. The application of alternative estimation methods to the same data set allows us to assess the robustness of the empirical findings.

## REVIEW OF RELEVANT LITERATURE

Empirically, the bulk of early literature using aggregate trade data from industrialized countries has generally reported an insignificant or weak impact of the exchange rate volatility on trade (Kenen, 1979; Thursby and Thursby, 1985). Evidence of negative impacts of exchange rate volatility on trade was only documented in the studies using more recent data of the flexible currency regime (Doroodian, 1999; Chowdhury, 1993). Still, some researchers fail to provide any significant impact of exchange rate volatility on trade flows (Abbott et al., 2001; Pattichis, 2003; Klaassen, 2004; Sigh, 2004). Baum et al. (2004) find 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 also found in Asseery and Peel (1991), McKenzie and Brooks (1997) and Daly (1998), among others.[5]

Turning to the literature on EMEs, which is the focus of the current study; we find that very limited number of studies has analyzed the relationship between exchange rate volatility and trade flows. A list of the research on EMEs between 1995 and 2010 is summarized in Table 1. In most cases, we find evidence of a negative relationship between exchange rate volatility on trade. Out of the eight studies in Table 1, four reported a negative and significant effect on exchange rate variability on trade. For instance, research by Doroodian (1999) and Doğanlar (2002) documented evidence of a contractionary impact of currency volatility on total export flows. Similarly, empirical studies by Rahmatsyah et al. (2002) and Siregar and Rajan (2004) provide support for the negative effects of currency volatility on bilateral trade flows of Thailand and Indonesia, respectively. Evidence of exchange rate volatility having a positive effect on trade has been reported by Kim and Lee (1996), Arize (1999) and Poon et al. (2005), while Sauer and Bohara (2001) documented some mix results from different specifications of panel regression model.

To note, two more recent studies (not reported in Table 1) by Hayakawa and Kimura (2009) and Chit et al. (2010) whom applied gravity model to a panel of East Asian countries also confirm the contractionary effect of exchange rate risk on bilateral export flows. A recent article by Hall et al. (2010) observed that the REX volatilities of EMEs (5.4%) are much lower than other developing countries that are not-EMEs (10.9%). According to Hall et al. (2010), financial market deregulation in the past two decades together with more disciplined macroeconomic policy has led to lower exchange rate uncertainty in the EMEs compared to other developing countries. Based on panel data from 1980-2006, they found that EMEs do not provide support of the hypothesis that exchange rate volatility has a negative impact on exports. Meanwhile, for the developing countries that are non-EMEs, the authors confirm a negative relationship between the two variables.

[5] Studies that used sectoral data generally obtained mix 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. However, in another stream of studies that uses the gravity approach, evidence of negative effects were consistently documented either using bilateral or multilateral, aggregate or sectoral data (Rose, 2000; Broda and Romalis, 2004; Clark et al., 2004; and Tenreyro, 2004). McKenzie (1999) provides an excellent literature survey on the issue of exchange rate volatility and trade.

In short, all these early studies are based on data of 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 shifting exchange rate regime and monetary policy adopted by the countries under study in the post crisis period. In our view, this issue needs further empirical verification by taking into account recent historical events to highlight the policy lessons of the 1997 Asian crisis.

## METHODOLOGY

Following Byrne et al. (2008) and others, we adopt the trade model developed by Armington (1969). Our empirical analysis is based on the following specifications:

$$\ln IM_t = \alpha + \beta_1 \ln Y_t^D + \beta_2 \ln(P_t^F - P_t^D) + \beta_3 \sigma_t^{ex} + \xi_t \tag{1}$$

$$\ln XP_t = \alpha + \beta_1 \ln Y_t^F + \beta_2 \ln(P_t^F - P_t^D) + \beta_3 \sigma_t^{ex} + e_t \tag{2}$$

where $\sigma_t^{ex}$ represent the conditional standard deviation of exchange rate and $\xi_t$ and $e_t$ are the usual residual terms. Based on prediction from theory, the elasticity (coefficients) for the income variable $\beta_1$, and the price variable $\beta_2$ are expected to be positive and negative, respectively. The elasticity for the currency volatility $\beta_3$ can be in either sign. In the case where all the variables in both equations are *I(1),* we could apply popular tests such as the Johansen and Juselius (1990) multivariate cointegration test, to investigate the existence of a long-run cointegrating relationship. The variables in our specifications are likely to be of different orders. For this purpose, we applied the bounds testing approach proposed by Pesaran et al. (2001). The major attraction of this the test is that the method can be applied to models irrespective of whether the regressors are $I(0)$ or $I(1)$, or mutually cointegrated. Thus, the technique 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:

$$\begin{aligned}\Delta \ln XP_t = a_0 + \sum_{i=1}^{p} b_i \Delta \ln XP_{t-i} + \sum_{i=1}^{p} b_i \Delta \ln Y_{t-i} + \sum_{i=1}^{p} b_i \Delta \ln(P_t^T - P_t^D)_{t-i} \\ + \sum_{i=1}^{p} b_i \Delta \ln \sigma_{t-i}^{ex} + \delta_1 \ln XP_{t-1} + \delta_2 \ln Y_{t-1} + \delta_3 \ln(P_t^T - P_t^D)_{t-1} + \delta_4 \ln \sigma_{t-1}^{ex} + \mu_t\end{aligned} \tag{3}$$

$$\begin{aligned}\Delta \ln IM_t = a_0 + \sum_{i=1}^{p} b_i \Delta \ln IM_{t-i} + \sum_{i=1}^{p} b_i \Delta \ln Y_{t-i} + \sum_{i=1}^{p} b_i \Delta \ln(P_t^T - P_t^D)_{t-i} \\ + \sum_{i=1}^{p} b_i \Delta \ln \sigma_{t-i}^{ex} + \delta_1 \ln IM_{t-1} + \delta_2 \ln Y_{t-1} + \delta_3 \ln(P_t^T - P_t^D)_{t-1} + \delta_4 \ln \sigma_{t-1}^{ex} + \mu_t\end{aligned} \tag{4}$$

The null hypothesis of non-existence of the long run relationship is tested using the F-test on ($H_0 : \delta_1 = \delta_2 = \delta_3 = \delta_4 = 0$) against ($H_0 : \delta_1 \neq 0, \delta_2 \neq 0, \delta_3 \neq 0, \delta_4 \neq 0$). 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. Upper bound assumes that all the variables in the ARDL model are $I(1)$ while 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, unit root test of stationarity is needed to authentic the order of integration of respective variables. If these series are found indeed 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. To this end, we form an unrestricted error correction model (ECM) to yield a more efficient estimate of the short-run coefficients (Stučka, 2004), given by:

$$\Delta y_t = -\phi(1,\hat{p})EC_{t-1} + \sum_{i=1}^{k}\beta_{i0}\Delta x_{it} + \delta^{'}\Delta w_t - \sum_{j=1}^{\hat{p}-1}\phi_j^{*}\Delta y_{t-j} - \sum_{i=1}^{k}\sum_{j=1}^{\hat{q}_i-1}\beta_{ij}^{*}\Delta x_{i,t-j} + \mu_t \qquad (5)$$

where $EC_t = y_t - \sum_{i=1}^{k}\hat{\theta}_i x_{it} - \hat{\psi}^{'} w_t$ , $y$ is the dependent variable, $x$ are the *i* number of regressors, $w$ are 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 normalizing the dynamic short run coefficients of each explanatory variable to one minus the short run coefficients of the dependent variable.

The analyses are based on monthly frequency data spanning from January 1990 to December 2008 for six East Asian EMEs, namely Indonesia, Malaysia, the Philippines, Singapore, Thailand and South Korea. All six countries have already embarked on financial market liberalization starting at this period. The countries were selected because exports are considered to be important for economic growth. We did not go beyond 2008 as the recent subprime crisis that started in 2009 is yet to be concluded. The time series data are import and export values, industrial production, Consumer Price Index (CPI), and nominal USD exchange rates. These are not a seasonally adjusted series. CPI is used to proxy domestic and foreign prices. All our data are downloaded from DataStream and the trade series (import and export) are extracted from the IMF Direction of Trade Statistics (DOTS).

Following the empirical literature, we focus on real exchange rate (REX hereafter).[6] We employed Nelson's (1990) Exponential Generalized Autoregressive Conditional Heteroskedasticity (EGARCH hereafter) to generate the series of conditional volatility of the REX. Then, the endogenous multiple structural break tests advocated by Bai and Perron (1998, 2003) (BP, hereafter) are utilized to detect possible breakpoints in the series.[7] The break tests reveal that two breaks are common in the REX series for all the countries. 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 is related to the 1997 Asian financial crisis, while the second break coincides with the adoption of inflation targeting in most of the crisis-affected countries. For Malaysia, the results reveal three structural breaks (January, 1997, February, 2000 and August, 2005) in the exchange rate volatility series, with the third break date occurring around the abolishment of hard peg and the lifting of the capital controls. We should authentically take into account these break points in the analysis that follows. In this study we find that when regime shifts in the condition variances are modeled, the persistence of the exchange rate volatility is much lower. In other words, accounting for the breaks weakens the persistence in volatility to some degree in the EGARCH model. To conserve space, the results of the Bai and Perron (BP) tests are not reported here but available upon request.  These breakpoints are then used to construct dummy variables for estimating exchange rate volatility.

The EGARCH model has several added advantages over the conventional GARCH model in capturing the REX volatility. The model is able to simultaneously accommodate asymmetric volatility and the leverage effect of bad news. A more important merit of this model is its ability to capture large shocks of any sign in financial series, which is particularly critical given the currency crash period in our sample. The model is specified in the following equations:

$$REX_t = \mu + \delta REX_{t-1} + \varepsilon_t + \varphi\varepsilon_{t-1} \text{ where } \varepsilon_t|\Omega_{t-1} \sim \text{GED}(r) \qquad (6a)$$

$$\log\sigma_t^2 = \omega + \sum_{i=1}^{p}\left(\alpha_i\left|\frac{\varepsilon_{t-i}}{\sqrt{\sigma_{t-i}}}\right| + \gamma_i\frac{\varepsilon_{t-i}}{\sqrt{\sigma_{t-i}}}\right) + \sum_{j=1}^{q}\beta_j\log\sigma_{t-j}^2 + \sum_{k=1}^{b}\eta_k Break_k \qquad (6b)$$

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[6] $REX_i = \ln[NEX_{i/j}(P_i/P_j)]$, where NEX the nominal exchange rate. The conditional standard deviation of REX generated from EGARCH is used as a proxy for exchange rate volatility. Previous literature in this area shows that there are no qualitative differences in using nominal or REX volatility; see for example McKenzie and Brooks (1997).

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

Equations (6a) and (6b) are the conditional mean and variance of the logarithmic of monthly REX, respectively. To get rid of linear dependency in 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 (6b) 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 (6a), $\mu$ is the intercept term; $\delta$ and $\varphi$ represents the magnitudes of the autoregressive and the moving average terms, respectively; and $\varepsilon_t$ is the idiosyncratic news. In the conditional variance equation (6b), $\sigma_{t-1}^2$ represents the lagged conditional variance of $\varepsilon_t$ while $\alpha$, $\beta$ and $\gamma$ 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, p.10). The dummy variable $Break_k$ with the parameters $\eta_k$ 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. The breaks dates on the volatility of the US dollar rates are determined using the BP procedure. Also, it is widely acknowledged that the standard 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. Bollerslev (1987) proposed student *t*-distribution to model fat-tailed property, while Nelson (1991) proposed a more general distribution for the error behavior—the Generalized error distribution (GED). We thus applied the GED as it is more flexible and nests several other distributions.

The summary of the descriptive statistics, correlation matrix and the EGARCH estimates for the exchange rate volatility are not reported here to converse space (but are available upon request). The mean of the exchange rates series are all positive, indicating that, on average, the Asian currencies are depreciating over the sample period. The conditional standard deviation estimated from model (6b) is a proxy for exchange rate volatility with a value smaller than zero. 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 behavior inherent in the exchange rate series. The use of GED can account for these non-normalities. The correlation matrix implies that all the series are not highly correlated so the problem of multicolinearity is not a worry. In the EGARCH estimates, the REX series show significant ARMA process, except for the Malaysian ringgit and Singapore dollar. In the variance equation, almost all the estimated parameters are highly significant, except for the asymmetric parameter, which is positive and significant only for the Philippine peso. Only the ringgit and Singapore dollar have a negative sign, which implies leverage effect, but they are not statistically significant.[8] This outcome is by and large due to the control of the structural break points in the variance equation.[9] Nearly all the estimates for the structural break parameter $\eta$ are positive and highly significant, implying that the exchange rate volatilities were significantly exaggerated during the endogenous detected structural distortions. The use of GED distribution obviously is justified as the parameter *r* value is tested to be significantly smaller than 2. Additionally, the diagnostic checking implies no major deficiency in the fitted model.

## EMPIRICAL RESULTS

The results of the bounds test are known to be sensitive to the order of lag length imposed on each of the first-differenced variable in each model (Bahmani-Oskooee and Goswami, 2003). Thus, to select the optimum number of lags on each variable, a maximum lag length of 12 was initially imposed in both Eqs.

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

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

(3) and (4). The dimensions of the parameter space were reduced to a final parsimonious set by sequentially eliminating statistically insignificant coefficients following Hendry's general-to-specific lag selection criterion. All variables with absolute *t*-values less than one were sequentially dropped. We do not report the bound test results here to save space. Generally, the computed *F*-statistics for all the estimated equations are greater than the upper bound of the critical values at 5% significant level or better in the export and import demand ARDL models. Clearly, this implies that there is a long-run relationship between exports (imports) and its determinants. The bounds test result is not reported here to conserve space.

We report the ECM estimates in Tables 2 and 3. Generally, the models perform well. The LM statistics indicate that autocorrelation is not a problem in any of the equations. There is no evidence of heteroscedaticity as indicate by the Engle's LM ARCH test. There is, however, evidence of non-normality in two equations (Singapore's exports and Thailand's imports). All of the equations, except for Singapore's import, easily pass the Ramsey's RESET test, indicating that the ECM is correctly specified. Note that the figures reported in parenthesis in Tables 5 and 6 are the Newey-West HAC standard error to address the Pagan's (1986) critique regarding the problems of generated variable. Given the evidence supporting the adequacy of the estimated model, we can make several observations regarding the results summarized in the two tables. First, in every country the majority of the short-run coefficient estimates are significant at 10% 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 expected. Third, the cumulative sum of the coefficient of the exchange rate volatility is negative in three out of six countries for the export equations, with values ranging from -0.033 to -0.015. The negative sign implies that an increase in exchange rate uncertainty reduces exports. Meanwhile, exchange rate volatility elasticity in the import equation turned out to be positive in all but one country (Korea). This outcome rejects the notion that exchange rate uncertainty adversely affects imports in most of the EMEs, at least in the short-run for the sample period considered. Finally, the coefficient estimate of the error correction term ($ECT_{t-1}$) is negative and statistically significant in all countries. The coefficient estimates range from 0.10 to 0.34 and 0.08 to 0.43 for exports and imports, respectively. 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 10% to 8% per month. Likewise, we observed that it is the imports that adjust to any shocks in the equation. Importantly, a significant negative $ECT_{t-1}$ reinforced the cointegrating relationship reported earlier. Also, we should mention that most of the test statistic from the CUSUM and CUSUM squares tests do not reject the null hypothesis of parameter stability at the 5% level.[10]

The ARDL bounds test confirms the presence of a long-run relationship between total exports (imports) and its determinants. For comparison, we also present the results of panel cointegration tests with four statistics suggested by Pedroni's (2001).[11] All in all, the computed statistics suggest the null of no cointegration is easily rejected for most of the estimations (see Table 4).[12] Next, the cointegrating relationship is estimated with the fully modified OLS estimator of Phillips and Hansen (1990). The FMOLS estimates of the individual equations are reported in Table 4. Several points merit comments. First, by focusing on the individual coefficients, we found that the income coefficient is positive and statistically significant at the 1% level. For the import function, the income elasticity ranged from 1.52 (Indonesia) to 3.05 (Malaysia). Similarly, we found that the magnitude of the income coefficient in the export equation is in the elastic range in all the countries under investigation (1.41 to 4.10). The relative price variable carries the expected negative sign and is significant at the usual significance levels in both sets of equations for

[10] To conserve space, the plots of both tests are not reported but available from the authors upon request.

[11] Pedroni (2004) has developed two sets of statistics to test the null of no cointegration. The first set of statistics is based on pooling the residuals along the 'within' dimension of the panel (the residuals of alternative hypothesis have common AR coefficient); whilst, the second set based on pooling the residuals along the 'between' dimension of the panel (assumed the residuals of alternative hypothesis have individual AR coefficient). Each of these statistics indeed has a comparative advantage in term of small sample size and power properties depending on the underlying DGP. From the Monte Carlo simulation, Pedroni (2004) found the group-rho statistics performed better in a very small panel because it is slightly undersized and empirically the most conservative of the tests. The other statistics tend to lie somewhere in between these two extremes, and they tend to have minor comparative advantages over different range of the sample size. (Pedroni, 2004, pp.616)

[12] We also estimate the long-run elasticities in both equations using the ARDL approach. The income variables are all significant at the 1% significance level and the elasticities are in the elastic range. For most of the other variables, the low *t*-values make it difficult to interpret. To conserve space, these results are not reported here but available upon request from the authors.

each of the countries. The magnitude of the coefficients is consistent with some studies mentioned earlier (e.g., Rahmatsyah et al., 2002).

Given our special interest is on the impact of exchange rate volatility on trade flows in the EMEs, it is important to discuss the role of this variable in some details. As shown in Table 4, magnitude and the sign exchange rate volatility in the long-run equation vary across countries. This finding reflects the underlying structure as well as the development of the capital markets in the countries under review. The hypothesis that exchange rate volatility has an adverse effect on trade flows is not rejected by the data in only three cases for the export equation, namely Korea (-0.17), Malaysia (-0.05) and Thailand (-0.06). For Indonesia and Singapore, the variable carries a positive sign but it is statistically insignificant, implying that exchange rate volatility has no economic impact in these two countries. The variable enters with a positive (0.18) sign and it is statistically significant at the 1% level for the Philippines. Unlike the other countries, the exchange rate risk has a positive effect on the Philippines' exports, a result that is consistent with the prediction by De Grauwe (1988) model. The mixed result in the export equation is also reflected in the import equations. For Indonesia, the Philippines, Malaysia and Thailand, our results do not show the negative and significant effect of exchange rate volatility on imports. On the other hand, it has adverse effects on Korea and Singapore imports. A striking feature of our results is that the estimated elasticities are small in comparison with most of the earlier studies (see Table 1).

Results using the group-mean FMOLS method of Pedroni (2001) suggest that on average the elasticity of exchange rate volatility for export and import functions are -0.09 ($t$-value=6.12) and -0.16 ($t$-value=6.58), respectively. Coefficient of the relative price has the expected negative sign and income elasticity is larger than unity. On the basis of the panel group estimates we may conclude that trade is sensitive to currency volatility albeit this response is small. To sum, the estimates of the short-run dynamics of the ECM indicate that exchange rate volatility has significant short-run negative effects on export demand in only three countries. We find that exchange rate volatility had negative long-run impact on the exports of Korea, Malaysia and Thailand, and a positive impact in the case of the Philippines. It has no significant (zero) effect on Indonesian and Singaporean exports. For imports, it carries a positive sign in the short-run dynamic model (except for Korea), but turned out to be negative and significant for two countries (Korea and Singapore) in the long-run equation. Finally, we added a dummy variable to capture the impact of the 1997 crisis and found that the coefficient was insignificant. This result (not reported) appears to indicate the adverse effect of uncertainty on trade is captured by the exchange rate volatility variable. How do our results compare with earlier studies reported in Table 1? Majority of the previous studies documented a significant contractionary effect of currency volatility. A consistent finding in the literature perhaps is that the currency volatility effects tend to be highly elastic when GARCH based volatility were used, except in Arize (1999); but when other volatility measure were used (predominantly MASD) the effect tends to be highly inelastic (smaller than 0.001 in many cases). Our estimates are all in the inelastic range, but at best provide mixed support for the hypothesis that it has negative effects of uncertainty on trade.[13]

## CONCLUSION

This study presents an empirical investigation of the hypothesis that exchange rate volatility has a negative impact on the exports and imports volume for six East Asian EMEs. Unlike the majority of past studies, we proxy exchange rate volatility using the EGARCH specification to reveal the short- and long- run relationships between exchange rate risk and trade. We find that the volatility model fits the data well and also confirms that exchange rate volatilities are affected by major economic events, including the 1997 Asian currency crisis. We show that the post-crisis era is characterized by higher exchange rate volatility in all but one country, Malaysia. This finding is consistent with the view that the exchange rate regime in these Asian countries is more flexible in the post-crisis era. Our findings show that the structural shifts in the volatility model have implications on the trade-exchange rate volatility relationship, and neglecting structural breaks may lead to misspecification of the conditional variance.

[13] For comparison with the other developing countries, Arize et al. (2008), for example, found that the exchange rate volatility gives a consistent negative impact for the Latin America countries ranging from -0.40 to -0.001. The long run income elasticity estimated reported by Arize et al.'s (2008) for eight Latin America countries ranged from 0.51 to 7.50, while the price ratio elasticity ranged from -2.29 to -0.29, with one positive case, 0.19 for Costa Rica.

The estimates of the long-run coefficients indicate that imports and exports are mostly income elastic, but price insensitive. We found that the statistical impact of exchange rate volatility is significant in some, but not all of the EMEs. Our results highlight that the REX volatility has a significant negative effect on exports in the short-run for Korea, Malaysia and Philippines. This negative link between volatility and exports is also observed in the long-run for Korea, Malaysia and Thailand. In contrast, exchange rate risk has an expansionary impact on the Philippine's export sector in the long-run. We found that exchange rate volatility has positive effect on the volume of imports in the short-run, except for one country (South Korea). Meanwhile, the impact of exchange rate risk on imports is at best mixed in the long run. It is reasonable to conclude that some evidence of a dampening effect of exchange rate volatility on trade flows is found in some, but not all, EMEs countries. The economic impact, however, is small in relation to the relative size of the other core variables—relative price and income. Recently, Hall et al. (2010) found that the impact of exchange rate volatility on exports in EMEs differs from other developing countries. Their results from a panel consisting of 11 EMEs do not support the hypothesis that exchange rate volatility had a negative effect on exports.

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 sectors. A further extension would consider a selection of important industries in these Asian countries. Such a research avenue would broaden our understanding of the relationship between trade performance and exchange rate uncertainty across sectors of the emerging Asian market economy.

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TABLE 1: Estimations from Existing Literature on the Same East Asia EMEs

| Paper | Trade Model, Sample and Period | Coefficient for Currency Volatility [volatility measure] |
|---|---|---|
| Kim and Lee (1996)<br>*Asian Economic Journal* | Model: ARCH<br>Sample: Korea<br>Period: 1980M1-1993M2 | **Real Aggregate Export Volume**<br>9.847**[ARCH-M] |
| Doroodian (1999)<br>*Journal of Asian Economics* | Model: ARIMA<br>Sample: Korea | **Aggregate Export**<br>-148.77* [GARCH] |
| | Sample: Malaysia<br>Period: 1973Q2-1996Q3 | -497.21* [GARCH] |
| Sauer and Bohara (2001)<br>*Review of International Economics* | Model: fixed and random effect panel model<br>Sample: a panel of 12 Asia developing countries, which include all the countries in the present paper except Singapore<br>Period: 1973-1993 (annual) | **Real Aggregate Export**<br>FE: 0.907, RE: 1.147 [ARCH]<br>FE: -1.037, RE: 0.282 [AR]<br>FE: -1.969, RE: 0.226 [Trend] |
| Arize (1999)<br>*The International Trade Journal* | Model: Johansen Cointegration<br>Sample: Singapore<br>Period: 1973Q2-1997Q1 | **Aggregate Trade**<br>0.06* [GARCH]<br>**Aggregate Export**<br>0.08* [GARCH] |
| Doğanlar (2002) | Model: Engle-Granger Cointegration | **Aggregate Export** |

| | | |
|---|---|---|
| *Applied Economics Letters* | Sample: South Korea<br>Period: 1980Q1-1996Q4 | -2.24 [MASD] |
| | Sample: Indonesia<br>Period: 1980Q1-1996Q4 | **Aggregate Export**<br>-1.08 [MASD] |
| | Sample: Malaysia<br>Period: 1980Q1-1994Q1 | **Aggregate Export**<br>-0.68 [MASD] |
| Rahmatsyah et al. (2002)<br>*Japan and the World Economy* | Model: ARDL<br>Sample: Thailand<br>Period: 1970Q1-1997Q2 | **Export to US** #<br>nil [GARCH]<br>(-0.0004)**,(-0.0003)*[MASD]<br>**Import from US** #<br>-0.0004** [GARCH]<br>-0.0004* [MASD] |
| Siregar and Rajan (2004)<br>*Journal of the Japanese and International Economies* | Model: Johansen Cointegration<br>Sample: Indonesia<br>Period: 1984Q1-1997Q2 | **Aggregate Export**<br>-150.257*[GARCH]<br>-0.003*** [MASD]<br>**Aggregate Import**<br>4.28 [GARCH]<br>-0.0002 [MASD] |
| | Sample: Indonesia<br>Period: 1984Q1-1997Q2 | **Export to Japan**<br>-29.62*** [GARCH]<br>-0.0009*** [MASD]<br>**Import from Japan**<br>-25.89*** [GARCH]<br>-0.0004*** [MASD] |
| Poon et al. (2005)<br>*ASEAN Economic Bulletin* | Model: Johansen Cointegration<br>Sample: Indonesia, Korea, Singapore, Thailand<br>Period: 1973Q2-2002Q2 | **Aggregate Export**<br>Indonesia: 0.42*** [MASD]<br>Korea: -0.23*** [MASD]<br>Singapore: -0.19*** [MASD]<br>Thailand: 0.09*** [MASD] |

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

TABLE 2: Error Correction Model for Export Model using ARMA(1,1)-EGARCH (1,1)

| | Indonesia | | Korea | | Malaysia | | Philippine | | Singapore | | Thailand | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| Intercept | -2.8079$^{a}$ | (0.5359) | -2.4760$^{a}$ | (0.5208) | -0.6690$^{a}$ | (0.1748) | -0.1546$^{a}$ | (0.0437) | -0.9676$^{a}$ | (0.3110) | -0.9678$^{a}$ | (0.1805) |
| Δln[XP(-1)] | -0.23370$^{a}$ | (0.0553) | -0.2435$^{a}$ | (0.0691) | -0.1420$^{a}$ | (0.0499) | -0.3808$^{a}$ | (0.0683) | -0.5749$^{a}$ | (0.0781) | -0.2515$^{a}$ | (0.0674) |
| Δln[XP(-2)] | 0.09356$^{b}$ | (0.0466) | 0.1312$^{b}$ | (0.0666) | | | | | -0.2369$^{a}$ | (0.0763) | 0.0931 | (0.0607) |
| Δln[XP(-3)] | | | 0.2158$^{a}$ | (0.0586) | | | | | 0.1236$^{b}$ | (0.0608) | 0.1503$^{a}$ | (0.0557) |
| Δln[XP(-4)] | | | | | | | -0.2621$^{a}$ | (0.0687) | | | -0.1853$^{a}$ | (0.0500) |
| Δln[XP(-5)] | | | | | | | -0.1550$^{b}$ | (0.0663) | -0.1704$^{a}$ | (0.0617) | | |
| Δln[IPUS)] | 1.7267$^{a}$ | (0.3760) | | | 1.7312$^{a}$ | (0.2773) | 2.3741$^{a}$ | (0.2866) | 0.9334$^{a}$ | (0.3017) | 1.0553$^{a}$ | (0.2402) |
| Δln[IPUS(-1)] | 1.2725$^{a}$ | (0.4210) | | | 1.9089$^{a}$ | (0.3028) | 2.3298$^{a}$ | (0.4429) | 1.5054$^{a}$ | (0.3610) | 1.8726$^{a}$ | (0.2889) |
| Δln[IPUS(-2)] | | | 0.7079$^{a}$ | (0.1940) | 1.8133$^{a}$ | (0.2551) | 0.9203$^{a}$ | (0.3428) | 1.8933$^{a}$ | (0.5130) | 1.7974$^{a}$ | (0.2302) |
| Δln[IPUS(-3)] | -1.4709$^{a}$ | (0.3471) | | | | | | | 0.7653$^{c}$ | (0.4126) | | |
| Δln[IPUS(-4)] | | | | | | | 1.1573$^{a}$ | (0.2664) | | | | |
| Δln[IPUS(-5)] | 1.4228$^{a}$ | (0.2472) | | | -0.4544$^{b}$ | (0.1824) | 1.2028$^{a}$ | (0.3113) | | | | |
| Δln[PD] | | | | | 1.3865 | (1.0320) | -1.4984$^{b}$ | (0.6610) | | | -0.9624 | (0.8273) |
| Δln[PD(-1)] | 1.0214$^{a}$ | (0.3673) | | | | | | | | | | |
| Δln[PD(-2)] | | | | | | | | | 2.4230$^{c}$ | (1.2490) | | |
| Δln[PD(-3)] | -0.9266$^{b}$ | (0.3810) | -3.0923$^{a}$ | (1.1160) | | | | | | | | |
| Δln[PD(-4)] | 0.8349$^{b}$ | (0.4051) | | | | | | | | | | |
| Δln[PD(-5)] | -1.4236$^{a}$ | (0.4990) | | | | | 1.5510$^{b}$ | (0.7744) | | | | |
| Δln[EXV2] | | | -0.0622 | (0.0405) | -0.0148$^{b}$ | (0.0073) | | | | | | |
| Δln[EXV(-1)] | | | | | | | | | 0.1458$^{b}$ | (0.0570) | | |
| Δln[EXV(-2)] | | | | | | | | | | | | |
| Δln[EXV(-3)] | 0.0556$^{c}$ | (0.0330) | 0.0552 | (0.0439) | | | | | | | 0.0450$^{c}$ | (0.0273) |
| Δln[EXV(-4)] | | | -0.0595$^{c}$ | (0.0355) | | | -0.0106 | (0.0130) | | | | |
| Δln[EXV(-5)] | | | | | | | -0.0222$^{c}$ | (0.0129) | | | | |
| ECT(-1) | -0.2306$^{a}$ | (0.0440) | -0.3358$^{a}$ | (0.0707) | -0.0959$^{a}$ | (0.0250) | -0.0689$^{a}$ | (0.0209) | -0.1290$^{a}$ | (0.0419) | -0.2124$^{a}$ | (0.0398) |
| *Diagnostic Checking* | | | | | | | | | | | | |
| LM(4) | 2.1396 | [0.0772] | 1.6828 | [0.1148] | 0.5079 | [0.7300] | 1.2508 | [0.2908] | 2.1538 | [0.0755] | 1.1759 | [0.3225] |
| ARCH (7) | 0.4618 | [0.8613] | 0.6731 | [0.6948] | 1.2337 | [0.2857] | 0.3873 | [0.9092] | 0.4128 | [0.8937] | 1.6103 | [0.1340] |
| Norm (2) | 1.0239 | [0.5993] | 0.2535 | [0.8809] | 5.2430 | [0.0727] | 3.5517 | [0.1693] | 8.5043$^{b}$ | [0.0142] | 4.4739 | [0.1068] |
| HET | 1.1211 | [0.3235] | 1.5715 | [0.0700] | 1.2651 | [0.2142] | 1.3628 | [0.1229] | 1.0637 | [0.3710] | 1.5594 | [0.0661] |
| RESET | 1.3189 | [0.2697] | 0.0046 | [0.9954] | 0.1361 | [0.8729] | 0.7812 | [0.4592] | 4.5700$^{b}$ | [0.0114] | 1.6276 | [0.1989] |

Notes: ($^{a}$), ($^{b}$), and ($^{c}$) denote significance at 1, 5 and 10% significant 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 3: Error Correction Model for Import Model using ARMA-EGARCH (1,1)

| | Indonesia | | Korea | | Malaysia | | Philippine | | Singapore | | Thailand | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| Intercept | -3.7549[a] | (0.9150) | -0.5663[a] | (0.2166) | -0.3320[a] | (0.1137) | -0.0970[c] | (0.0523) | -0.9988[a] | (0.2081) | 0.3283[a] | (0.0606) |
| Δln[IM(-1)] | -0.3478[a] | (0.0508) | -0.4667[a] | (0.0628) | -0.4939[a] | (0.0518) | -0.5480[a] | (0.0780) | -0.2920[a] | (0.0696) | -0.2408[a] | (0.0722) |
| Δln[IM(-2)] | -0.1896[a] | (0.0535) | -0.3474[a] | (0.0772) | -0.1931[a] | (0.0580) | -0.2239[a] | (0.0518) | -0.2178[a] | (0.0576) | | |
| Δln[IM(-3)] | | | -0.2087[a] | (0.0604) | | | -0.1626[a] | (0.0597) | | | | |
| Δln[IM(-4)] | | | | | -0.1335[a] | (0.0511) | -0.1014[b] | (0.0497) | | | | |
| Δln[IM(-5)] | | | | | | | | | | | 0.1030[b] | (0.0486) |
| Δln[IPUS)] | -1.8482[a] | (0.5972) | 0.9286[a] | (0.3332) | | | 0.9716[a] | (0.2879) | 1.9716[a] | (0.3452) | | |
| Δln[IPUS(-1)] | -1.7348[a] | (0.6362) | | | 1.6480[a] | (0.4858) | 0.9653[a] | (0.3393) | | | | |
| Δln[IPUS(-2)] | | | 0.7210[b] | (0.2878) | 2.0282[a] | (0.5033) | 1.6610[a] | (0.3318) | | | 0.9553[c] | (0.4834) |
| Δln[IPUS(-3)] | | | | | | | 0.5871[b] | (0.2950) | -1.2148[a] | (0.3914) | | |
| Δln[IPUS(-4)] | | | | | | | | | | | | |
| Δln[IPUS(-5)] | -1.8184[a] | (0.6341) | -1.3042[a] | (0.3599) | | | | | -1.1251[a] | (0.4275) | | |
| Δln[PD] | 0.7740 | (0.7849) | 2.9631[c] | (1.5990) | | | 2.1727[a] | (0.7869) | | | | |
| Δln[PD(-1)] | | | 4.4085[b] | (0.0102) | -3.0954[c] | (1.6590) | | | -3.5010[b] | (1.7230) | | |
| Δln[PD(-2)] | | | | | | | | | | | | |
| Δln[PD(-3)] | | | | | | | | | -3.3868[b] | (1.6250) | -6.0771[b] | (2.3810) |
| Δln[PD(-4)] | | | | | -2.7072[c] | (1.4930) | 0.5396 | (0.7532) | | | | |
| Δln[PD(-5)] | | | 3.7507[a] | (1.2230) | | | | | | | | |
| Δln[EXV] | | | | | | | 0.0543[a] | (0.0185) | | | | |
| Δln[EXV(-1)] | 0.2364[a] | (0.0604) | | | 0.0352[a] | (0.0126) | -0.0243 | (0.0165) | | | 0.3292[a] | (0.0751) |
| Δln[EXV(-2)] | | | | | | | | | 0.1607[c] | (0.0872) | | |
| Δln[EXV(-3)] | | | -0.0906[c] | (0.0485) | | | | | | | | |
| Δln[EXV(-4)] | | | | | | | | | 0.1356 | (0.0884) | 0.1779[a] | (0.0679) |
| Δln[EXV(-5)] | | | | | | | | | | | 0.2039[b] | (0.0900) |
| ECT(-1) | -0.2628[a] | (0.0636) | -0.1401[a] | (0.0513) | -0.1228[a] | (0.0430) | -0.0777[b] | (0.0380) | -0.2805[a] | (0.0573) | -0.4324[a] | (0.0749) |
| *Diagnostic Checking* | | | | | | | | | | | | |
| LM(4) | 1.7115 | [0.1082] | 1.2710 | [0.2827] | 0.4777 | [0.7521] | 1.8439 | [0.1219] | 1.7071 | [0.1497] | 1.0700 | [0.3724] |
| ARCH (7) | 0.6531 | [0.7116] | 0.4922 | [0.8397] | 1.2865 | [0.2584] | 0.9905 | [0.4392] | 0.7849 | [0.6007] | 0.7253 | [0.6507] |
| Norm (2) | 1.8822 | [0.3902] | 3.0846 | [0.2139] | 1.7633 | [0.4141] | 4.0759 | [0.1303] | 1.4832 | [0.4763] | 29.821[a] | [0.0000] |
| HET | 1.3100 | [0.1939] | 1.0993 | [0.3502] | 0.7414 | [0.7653] | 1.0799 | [0.3444] | 1.1249 | [0.2760] | 0.5816 | [0.8957] |
| RESET | 4.2316[b] | [0.0158] | 1.7609 | [0.1745] | 0.1849 | [0.8313] | 0.2559 | [0.7744] | 1.0490 | [0.3521] | 2.7928 | [0.0635] |

Notes: ([a]), ([b]), and ([c]) denote significance at 1, 5 and 10% significant 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 4: FMOLS Long Run Estimate

| | Panel A: Export (to US) | | | Panel B: Import (from US) | | |
|---|---|---|---|---|---|---|
| | IP(US) | REX | EXV | IP(US) | REX | EXV |
| Countries specific coefficients | | | | | | |
| Indonesia | 2.77[a] | -0.37[b] | 0.03 | 1.52[a] | -0.92[a] | 0.01 |
| | [16.82] | [-2.52] | [0.68] | [5.61] | [-3.81] | [0.07] |
| Korea | 2.67[a] | -0.02 | -0.17[a] | 2.06[a] | -0.91[a] | -0.07[c] |
| | [23.69] | [-0.17] | [-5.25] | [15.96] | [-6.21] | [-1.79] |
| Malaysia | 4.10[a] | -1.16[a] | -0.05[b] | 3.05[a] | -1.16[a] | -0.02 |
| | [19.29] | [-5.48] | [-2.09] | [14.20] | [-5.42] | [-1.02] |
| Philippine | 2.25[a] | -0.23 | 0.18[a] | 2.67[a] | -0.50[a] | 0.03 |
| | [11.21] | [-1.22] | [2.88] | [16.41] | [-3.28] | [0.68] |
| Singapore | 1.41[a] | -1.49[a] | 0.05 | 2.52[a] | -1.20[a] | -0.18[a] |
| | [10.54] | [-8.39] | [0.70] | [21.45] | [-7.67] | [-3.09] |
| Thailand | 3.00[a] | -0.54[a] | -0.06[b] | 2.49[a] | -1.11[a] | -0.07 |
| | [22.55] | [-4.46] | [-2.25] | [11.42] | [-5.59] | [-1.50] |
| Panel Group | 2.70[a] | -0.64[a] | -0.01[b] | 2.38[a] | -0.97[a] | -0.05[a] |
| | [42.50] | [-9.08] | [-2.17] | [34.72] | [-13.06] | [-2.71] |
| Pedroni Cointegration tests | | | | | | |
| *Common (within-dimension)* | | | | | | |
| Panel rho | -17.327[a] | | | -17.320[a] | | |
| Panel PP | -10.670[a] | | | -11.029[a] | | |
| *Individual (between-dimension)* | | | | | | |
| Group rho | -22.497[a] | | | -22.283[a] | | |
| Goup PP | -13.878[a] | | | -13.609[a] | | |

Notes: ([a]), ([b]) and ([c]) indicate significant at 1, 5 and 10% significance levels, respectively. The values in [ ] denote the t-statistic. IP(US)=US industry production, PD=relative price, REX=real exchange rate and EXV=exchange rate volatility with breaks by ARMA(1,1)-EGARCH(1,1). Critical values for the four panel cointegration statistics ($H_0$: No cointegration) are from standard normal distribution. See Pedroni (2001) for details of computing the test statistics and their critical value.