The effects of exchange rate volatility on exports: Some new evidence for regional ASEAN countries

By

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ABSTRACT

The paper investigates empirically the impact of bilateral exchange rate volatility on the export flow of five regional ASEAN countries, namely Malaysia, Singapore, the Philippines, Indonesia and Thailand, to the United States, over the period January, 1990 to December, 2010. Estimates of the cointegration relations are obtained using methods proposed by Johansen and Juselius (1990). Furthermore, the short-run and long-run dynamic relationships between the variables are obtained for each country utilizing error correction modelling. In general, the real bilateral exchange rate volatility has a significant impact on exports at least for all the countries considered in our sample, and the impact overall is negative except for Indonesia.

1 INTRODUCTION

Exchange rate volatility is a source of concern as currency values partially determine the price paid or received for output/goods and, consequently, this affects the profits and welfare of producers and consumers (Akhtar and Spence, 1984)[1]. As a result, exchange rate volatility can affect the volume of goods traded internationally by making prices and profits indeterminate.

Theory suggests that exchange rate volatility may effects exports negatively or positively (Doyle, 2001 and Baak, 2004), based on the role played by the agents in the market. If economic agents are moderately risk averse the impacts of exchange rate volatility to exports will be negative (Cushman, 1983 and Koray and Lastrapes, 1989). This negative impact may come directly through uncertainty and adjustment costs, and indirectly through its effect on allocation of resources and government policies (Cote, 1994). Secru and Uppal (2000) showed the theoretical possibility of both positive and negative relationships, and Baccheta and Wincoop (2000) illustrated a theoretical model exhibiting no relationship between these variables. Numerous studies showing the high degree of volatility of exchange rate movements have led policy makers and researchers to investigate the nature and extent of the impact of such movements on volume of trade, both exports and imports.

[1] See Choudhry (2005) for citations of papers that explain the mechanism by which exchange rate volatility affects trade internationally.

In particular, previous empirical researchers have examined how the effect of both real and nominal exchange rate volatility on international trade depends on various factors. The overall evidence is best characterized as mixed as the results are sensitive to the choices of sample period, model specification, proxies for exchange rate volatility, and countries considered (developed, developing, Asia, ASEAN, etc). Yet, the relationship is still important enough to be explored especially for the principal ASEAN countries namely, Singapore, Malaysia, Thailand, Philippines, and Indonesia, due to various macroeconomic events, for instance the Asian financial crisis in 1997/1998. Due to these events the relationship between its major trading partners like the United States, is of interest. Moreover, for most of these countries export activity have been one of the major engines of economic growth. Thus in the light of international trade, the purpose of this study is to investigate the impact of five ASEAN countries exchange rate volatility on exports to the United States.

The impact of the exchange rate volatility on trade volume (in this case, exports) has been investigated in a significant number of studies, both theoretically and empirically. A detailed literature survey on the effects of exchange rate volatility on trade has been outlined by, among others, Cote (1994), McKenzie (1999) Clark, Tamirisa and Wei (2004) and Ozturk (2006). According to these surveys, exchange rate volatility can encourage the export volume through various factors[2]. From these factors the ultimate relationship between exchange rate volatility and the export volume can be categorize into three types of relationship as follows;

Type 1: The exchange rate volatility affects exports negatively (significant or not significant)
Type 2: The exchange rate volatility affects exports positively (significant or not significant)
Type 3: There are no relationships between these variables.

Broad discussions of this topic are covered by previous researchers, for example, Hooper and Kohlhagen (1978), Gotur (1985), Brada and Mendez (1988), Peree and Steinherr (1989), Klein (1990), Feenstra and Kendall (1991), Hook and Boon (2000), Doyle (2001), Baak (2004), among others. For more recent studies, see Arize et al. (2005), Lee and Saucier (2005), Baak et al. (2007), Chit et al. (2008), Aize (2008) and Baak (2009). However, the issue was rarely investigated for the exports of ASEAN countries. So far, only a small number of studies e.g. Arize et al. (2000), Baum et al. (2001), Doganlar (2002), Bahmani-Oskooee and Goswami (2004), Baak et al. (2007) have focus on or included ASEAN countries in their analyses.

[2] From the types of exchange rate (nominal or real exchange rate), types of countries (developing or developed), group of countries (ASEAN, Asian, EU, OECD, etc).

Some empirical evidence from these survey such as Akhtar and Hilton (1984), Cushman (1986), Peree and Steinherr (1989), Bini-Smaghi (1991), Savvides (1992), Chowdhury (1993), Hook and Boon (2000), Baak (2004), Arize et al. (2005), Lee and Saucier (2005), Baak et al. (2007), Chit et al. (2008), Arize et al. (2008) and Baak (2009) supports an increase in exchange rate risk having negative effect on the volume of exports.

In contrast, the evidence from other researchers such as Sercu and Vanhulle (1992), Baccheta et al. (2000), Aristotelous (2001), Bahmani et al. (1993), Gagnon (1993), Doyle (2001) and Bredin et al. (2003) demonstrated that the effect between exchange rates volatility and trade is positive or ambiguous. Following the work of Das (2003), Kasman and Kasman (2006), Arize et al. (2005), Baak (2007, 2008) and Arize et al. (2008) among others, examines the long run and the short run relationship between exchange rate volatility and exports by implementing cointegration tests and Granger causality tests in vector error correction modelling in their study.

Given these contradictory theoretical predictions and empirical findings, this study presents additional empirical evidence about the influence of exchange rate volatility on exports. Specifically, what we examine differs from that reported to date in a few major ways. Firstly, this paper uses data from a period of time that has not yet been used to investigate this question. The influence of exchange rate volatility on ASEAN and the United States bilateral export demand function from January, 1990 to December, 2010. A second distinguishing feature of this paper pertains to the measurement of exchange rate variability. Here, common[3] measure of exchange rate volatility is employed – the generalized autoregressive conditional Heteroscedastic (GARCH 1,1) model[4] of Bollerslev (1986). Kroner and Lastrapes (1993), Caporate and Doroodian (1994, Lee(1999) and Choudhry (2005), also apply the GARCH model to estimate the volatility of exchange rate. Finally, this study imposes an Asian Financial Crisis Dummy in the model from July, 1997 to December, 1999 in order to capture the structure break in the model. By giving this setting, the present study aims to determine whether the bilateral exchange rate volatility between ASEAN countries and its trading partner negatively affects the exports of the ASEAN countries.

[3] Here, the uncommon way in measuring the volatility is known as standard deviation of percentage changes in the exchange rate (De Grauwe and Bellefroid (1986) ; De Grauwe (1987, 1988); and Baak (2007)).

[4] According to Jansen (1989), the unconditional (uncommon) measure of volatility lacks a parametric model for the time varying variance of a time series. Therefore, referring to Arize (1995), the exchange rate volatility may be modeled by the Autoregressive Conditional Heteroscedastic (ARCH) model of Engle (1982). Furthermore, in this paper the conditional variance of the first difference of the log of the exchange rate is applied as volatility. The conditional variance is estimated by means of the Generalized Autoregressive Conditional Heteroscedastic (GARCH) model of order (1,1).

The rest of this paper is structured as follows. After this introduction section, the next section covers for the model specification and data used in this study. In section three, there is an overview of the econometrics approach used to estimate the parameters of the models. Section four presents the empirical finding and the final section summarizes and concludes the paper.

2 MODEL SPECIFICATION AND DATA

This research investigates the long run and short run relationship between exchange rate volatility and exports by performing Granger causality test in the vector error correction (VECM) framework, as in the studies of Baak (2001, 2007, 2008) and Arize Osang and Slottje (1999, 2000). Following the typical specification of others, and with additional specification as stated earlier in the introduction, the long run equilibrium relationship between exports and other economic variables in this study is examined based upon the following export demand equation:

$$\ln A_{ijt} = \alpha_0 + \alpha_1 \ln G_{jt} + \alpha_2 \ln P_{ijt} + \alpha_3 \ln \sigma_{ijt} + \alpha_4 CD_{ijt} + v_{ijt}$$

Here, A_{ijt} denotes as real exports from a country *i* (for example, Malaysia or Singapore) to a country *j* (the United States); G_{jt} is the GDP of an importing country, *j* ; P_{ijt} is the real bilateral exchange rate, reflecting the price competitiveness; σ_{ijt} is the volatility of the bilateral real exchange rates; CD_{ijt} is representing the crisis dummy due to the Asian financial crisis in mid 1997 to the end of 1999; finally v_{ijt} denotes as a disturbance term. All variables are in natural logarithms and the subscript *t* indicates the time period.

In the equation, thevariable G_{ijt} is used as a proxy for the level of economic activity in the importing country, in this case is the United States. It is expected that, the higher the economic activity in the importing country, the higher the demand for exports (Cote, 1994). Therefore, the value for α_1 is expected to be positive. Since the higher real exchange rate implies a lower relative price, the value for α_2 is also expected to be positive (Arize et al. (2000). As stated earlier, exchange rate volatility may effects trade negatively or positively. However, if the economic agents are moderately risk averse, as De Grauwe (1988) shows, it is generally expected that the impact of exchange rate volatility is negative. Thus, in this study because of the assumption of the economic agents is avoiding the risk, so the value for α_3 will be negative. Finally, a dummy variable (CD_{ijt}) is included in the model to represent the

Asian financial crisis in 1997/1998. In this case, CD=1 for the period from July, 1997 to December, 1999, and zero otherwise.

2.1 DATA DEFINITIONS

In estimating this model each variable definition follows the guidance proposed by Baak et al. (2007).

Crisis Dummy *(CD_{ijt})*

In order to capture the structure break of the Asian financial crisis from July 1997 to December, 1999, we introduce the crisis dummy variable denoted as CD_{ijt} above. There will be always a significant improvement in the stochastic properties of the VAR model is obtained by adding dummy/dummies to capture this historical episodes (Baharumshah, 2009). Thus, it is treated as an exogenous variable in the system. This study assumed there will be significant impact of crisis period (CD_{ijt}) on exports.

Economic Activity *(Gjt)*

As stated in Baak et al. (2007, 2008), the real GDP of the importing country j (in this case is the United States) is commonly used as a proxy measure for economic activity. However, due to the data availability for monthly data, in this study we prefer to use the production index as a proxy to economic activity. The same measurement has been used by the previous literature such as Chou and Shih (1998), (2000), Chou (2000), Baum (2001), among others.

Real Exports *(A_{ijt})*

Real exports from country *i* to country *j* can be expressed as follows,

$$A_{ijt} = \ln\left(\frac{a_{ijt}}{auv_{it}} \times 100\right)$$

where A_{ijt} denotes as the log value of the real total exports of country *i* to country *j*. While, a_{ijt} can be specify as a monthly nominal exports of country *i* to *j* and the auv_{it} is the export unit value index of country *i*.

Exchange Rate Volatility *(σ_{ijt})*

For the exchange rate volatility calculation, this study we apply one of the most popular volatility specifications, namely GARCH (1,1), as follows:

$$\sigma^2{}_{ijt} = \alpha_0 + \alpha_1 \varepsilon^2_{ijt-1} + \sigma^2_{ijt-1}$$

Real Bilateral Exchange Rate *(P_{ijt})*

The bilateral trade between two countries depends upon, among other factors, exchange rates and the relative price level of the two trading partners. Hence, the real exchange rates are included in the export equations of this paper and are computed as follows;

$$P_{ijt} = \ln\left(EX_{ijt} \times \left[\frac{CPI_{jt}}{CPI_{it}} \right] \right)$$

Here, P_{ijt} denotes the real monthly exchange rate in natural logarithm scale, while, EX_{ijt} is the nominal monthly exchange rate. The CPI_{it} and CPI_{jt} symbolize the monthly data for consumer price indexes of an exporting country *i* and an importing country *j*, respectively.

2.2 DATA SOURCES

This study uses monthly data covered from January, 1990 to December, 2010. Overall this study uses over 210 observations. The data such as consumer price indices (CPI), export unit value indices, production indices of the United State, and nominal exchange rates were obtained from the International Financial Statistics (IFS) of the International Monetary Fund (IMF). The data for exports from each ASEAN country to the United States were collected from the Direction of Trade Statistics (DOTS) of the IMF.

3 METHODOLOGY

This section provides methodology framework for the study. Arize et al. (2008) suggest that the multivariate analysis may lead a precise analysis in order to capture the long run and short run relationships between the variables. Therefore, in order to capture these issues, we apply both cointegration analysis and error correction model. However, before we go further, the beginning of a multivariate analysis for time series data lies in the univariate unit root tests. In this study, we perform the unit root tests proposed by Dickey and Fuller (1979).

3.1 AUGMENTED DICKEY FULLER (ADF) UNIT ROOT TESTS

Before we proceed to deeper analysis, we first evaluate the univariate and multivariate properties of the data set. We begin our analysis by estimating the unit root test proposes by Dickey (1976), Fuller (1976), and Dickey (1979). The well known Augmented Dickey Fuller tests use a parametric autoregression to approximate the ARMA structure of the errors in the test regression.

The unit root tests described above are valid if the time series, for example growth *(g_t)* is well characterized by an AR (1) with white noise errors. Many financial time series, however, have a more complicated dynamic structure than is captured by a simple AR (1) model. Said and Dickey (1984) augment the basic autoregressive unit root test to accommodate general

ARMA(p, q) models with unknown orders and their test is referred to as the augmented Dickey- Fuller (ADF) test. The ADF test tests the null hypothesis that a time series g_t is *I(1)* against the alternative that it is *I(0)*, assuming that the dynamics in the data have an ARMA structure.

The Augmented Dickey Fuller model structure is as follows. Consider a general AR (p) process given by,

$$g_t = \eta + \varphi_1 g_{t-1} + \varphi_2 g_{t-2} + ... + \varphi_p g_{t-p} + \varepsilon_t \tag{1}$$

If this is the process generating the data but an AR (1) model is fitted, say

$$g_t = \eta + \varphi_1 g_{t-1} + \upsilon_t \tag{2}$$

Then,

$$\upsilon_t = \varphi_2 g_{t-2} + ... + \varphi_p g_{t-p} + \varepsilon_t \tag{3}$$

And the autocorrelations of vt and v_{t-k} for $k > 1$; will be nonzero, because of the presence of the lagged ‘g’ terms. Thus an indication of whether it is appropriate to fit an AR (1) model can be aided by considering the autocorrelations of the residual from the fitted models. To illustrate how the DF test can be extended to autoregressive processes of order greater than 1, consider the simple AR (2) process below,

$$g_t = \eta + \varphi_1 g_{t-1} + \varphi_2 g_{t-2} + \varepsilon_t \tag{4}$$

Then notice that this is the same as:

$$g_t = \eta + (\varphi_1 + \varphi_2) g_{t-1} - \varphi_2 (g_{t-1} - g_{t-2}) + \varepsilon_t \tag{5}$$

And subtracting g_{t-1} from both sides gives:

$$\Delta g_t = \eta + \beta g_{t-1} + \alpha_1 g_{t-1} + \varepsilon_t \tag{6}$$

Where the following have been defined:

$$g_t = \varphi_1 + \varphi_2 - 1 \tag{7}$$

And

$$\alpha_1 = 1 \varphi_2 \tag{8}$$

Therefore to perform a Unit Root test on an AR (p) model the following regression should be estimated:

$$\Delta g_t = \eta + \beta g_{t-1} - \sum_{j=1}^{p} \alpha_j \Delta g_{t-j} + \varepsilon_t \tag{9}$$

The Standard Dickey-Fuller model has been 'augmented' by Δg_{t-j}. In this case the regression model and the *'t'* test are referred as the ADF test.

3.2 JOHANSEN JUSELIUS COINTEGRATION AND ERROR CORRECTION MODELLING (ECM)

We next proceed to use cointegration techniques pioneered by Engle and Granger (1987). Hendry (1986) and Granger (1986) made a significant contribution towards testing Granger causality. Two or more variables are said to be cointegrated (i.e.: they exhibit long run equilibrium relationship(s)), if they share common trends. According to this technique pioneer ed by Granger (1969) and Sims (1972), if two variables are cointegrated, the finding of no causality in either direction one of the possibilities with the standard tests, is ruled out. As long as the two variables have common trends, causality must exist in at least one direction, either unidirectional or bidirectional. Evidence of cointegration among variables also rules out the possibility of the estimated relationship being 'spurious'. Although cointegration indicates the presence or absence of Granger causality, it does not indicate the direction of causality between variables. This direction of the Granger (or temporal) causality can be detected through the vector error correction model derived from the long run cointegration vectors.

Moreover, this cointegration test is vital due to its ability in testing the relationship between non-stationary time series variables. If two or more series have a unit root, that is I(1), but a linear combination of them is stationary, I(0), then the series are said to be cointegrated. There are a few methods can be used to test for cointegration. One of most popular tests is by Johansen (1988) and Johansen (1990), since this particular method is often claimed to be superior to the regression-based Engle and Granger (1987) procedure. Lag truncation under this method proposed by Vahid and Engle (1989) is applied.

Suppose the vector of n-variables, $Z_t = (Z_{it}, Z_{2t}, Z_{3t}, ..., \ Z_{nt})$, is generated by the k^{th} order vector autoregressive process with Gaussian errors;

$$Z_t = \Phi + \phi_1 Z_{t-1} + \phi_2 Z_{t-2} + \phi_3 Z_{t-3} + ... + \phi_4 Z_{t-k} + \varepsilon_t \tag{10}$$

And,

$t = 1,2,3,4,5,.......T.$

where Z_t is a ($p \times 1$) vector of stochastic variables, and, $\varepsilon_1,......, \varepsilon_T$ are i.i.d. with normal distribution $\left(0, \sigma^2\right)$, mean zero and constant in variance. Since we want to distinguish between stationary by linear combination and by differencing this process may be written in error correction modelling or vector error correction (VEC) modelling form as equation below;

$$\Delta Z_t = \Phi + \Gamma_1 \Delta Z_{t-1} + \Gamma_2 \Delta Z_{t-2} + \Gamma_3 \Delta Z_{t-3} + ... + \Gamma_{k-1} \Delta Z_{t-k+1} + \Pi ECT_{t-1} + \varepsilon_t \tag{11}$$

And,

$t = 1,2,3,4,5,.......T.$

Based on the equation above, the matrix Π contains information about the long run relationship between the variables in the vector. Information about the number of cointegrating vectors is found in the rank of Π. In other words, the rank of Π determines how many linear combinations of Z_t vector are stationary. If the $(p \times p)$ matrix Π has rank r equal to zero, then all elements of Zt are non-stationary. Thus, there are no cointegration relationships between the variables. If *Π* is of full rank *r* = *p*, then all elements of Z_t are stationary. Thus, any combination of the variables results in a stationary series is cointegrated. In the intermediate case, when *r* < *p*, there are r non-zero cointegrating vectors among the elements of Z_t and *p-r* common stochastic trends. If a non-zero relationship is indicated by the test, a stationary long-run relationship is implied. Also, in equation (11), the notation 'ECT' indicates the error correction term that generated from the cointegration equation.

In the case where *0 < r < p*, Π can be factored as $\alpha\beta'$ (or $\Pi = \alpha\beta'$) where α and β are both ($p \times r$) matrices. The matrix α contains the adjustment parameters while β is called the cointegrating matrix and has the property that $\beta Z_t \sim I(0)$, where I(0) indicates integrated of order zero. Thus we can interpret the relations of βZ_t as the stationary relations among potentially non-stationary variable that is, as cointegrating relations. Johansen (1990) developed a maximum likelihood estimation procedure for Φ, Γ, α and β. This method also provides tests for the number of cointegration vectors; the λ_{trace} and $\lambda_{\max}$ formulation is as follows;

$$\lambda_{trace} = -T \sum_{i=r+1}^{n} \ln(1-\lambda_i) \tag{12}$$

Where T is the sample size and λ_{r+1},........, λ_{r+i} is the ordered $p-r$ smallest eigenvalues. The λ_{trace} statistic tests the null hypothesis that there are not at most r cointegrating vectors against a general alternative. However, the second statistic, the $\lambda_{\max}$ statistic, tests the null hypothesis that there are r cointegrating vectors against the alternative that there are $r+1$ cointegrating vectors. This statistic is written as;

$$\lambda_{\max} = -T \ln\left(1-\lambda_{r+1}\right) \tag{13}$$

Here in equation above, λ_{r+1} is an estimated eigenvalue. The critical values for the λ_{trace} and $\lambda_{\max}$ statistics are provided in Johansen (1990) and Osterwald-Lenum (1992).

4 EMPIRICAL RESULTS

In this section we discuss empirical results for this study. In the beginning, we discuss all the results in general, starting from ADF unit root tests to the Johansen cointegration tests, vector error correction modelling and regression equations of each country. Detailed discussion is provided at the end of this section.

4.1 THE AUGMENTED DICKEY FULLER (ADF) TESTS RESULT

The univariate Augmented Dickey Fuller (ADF) unit root test conducted in the all systems under this study concludes that all the data are I(1) processes. The unit root tests are employed to investigate the stationarity of the macroeconomic series at levels and then at first difference of each series.

To ensure the disturbances in all these equations are white noise, a sufficient number of lagged dependent variables have been included in the estimated regression[5]. The results of the tests, both at levels and at first differencing, are reported in Table 1, both with and without a time trend variable in the regression. Based on Table 1, the t-test statistics for all series from ADF tests are statistically insignificant to reject the null hypothesis of non-stationary at 0.01 significant levels. This result indicates that this series are non-stationary in their levels form. As the tests fail to reject the null hypothesis of unit roots in their level form in the autoregressive representation of each variable, thus, they are all not I(0). Therefore, these variables contain a unit root process or they share a common stochastic movement.

Thus, the tests continue to the first differencing stages. When the ADF test is conducted for first differences of each variable, the null hypothesis of non-stationary is easily rejected at 0.01 significant levels as shown in Table 1. Obviously, this result is consistent with some of the previous studies, e.g, Das (2003), Kasman and Kasman (2005), Arize et al. (2005), Baak (2007, 2008) and Augustine et al. (2008) among others. As claimed by Nelson and Plosser (1982), most of the macroeconomics and financial series are expected to contain a unit root and thus are integrated of order one, I(1). Therefore, this study concludes that the series are integrated of order one, and a higher order of differencing is not required.

[5] For the optimal lag length estimator, this study employs the Akaike Information Criteria (AIC).

Table 1: The Augmented Dickey Fuller (ADF) unit root tests results

Data Series	At Level		At first difference	
Singapore				
Data Series	**without time**	**with time**	**without time**	**with time**
AS_ijt	-2.537878 (4)	-2.520144 (4)	-8.964511 (4)a	-9.010142(4)a
GS_jt	-1.511249 (4)	-1.054364 (4)	-6.952419 (4)a	-7.041146 (4)a
PS_ijt	-0.897839 (4)	-0.723620 (4)	-6.237290 (4)a	-6.242019 (4)a
GARCH_ijt	-1.918365 (4)	-2.006650 (4)	-6.795549 (4)	-6.800304 (4)a
Malaysia				
Data Series	**without time**	**with time**	**without time**	**with time**
AM_ijt	-2.392370 (2)	-1.919868 (4)	-7.933085 (4)a	-8.275650 (4)a
GM_jt	-1.511244 (4)	-1.054364 (4)	-6.952419 (4)a	-7.041146 (4)a
PM_ijt	-1.511563 (4)	-1.322453 (4)	-6.261529 (4)a	-6.299796 (4)a
GARCH_ijt	-1.536856 (4)	-1.996264 (4)	-5.777183 (4)a	-5.768497 (4)a
Thailand				
Data Series	**without time**	**with time**	**without time**	**with time**
AT_ijt	-2.334833 (4)	-3.132127 (4)	-8.729152 (4)a	-8.79133 (4)a
GT_jt	-1.511249 (4)	-1.054364 (4)	-6.952419 (4)a	-7.04114 (4)a
PT_ijt	-1.495946 (4)	-1.180035 (4)	-5.915259 (4)a	-5.96630 (4)a
GARCH_ijt	-2.384251 (5)	-2.984798 (4)	-10.11501 (4)a	-10.1146 (4)a
Philippines				
Data Series	**without time**	**with time**	**without time**	**with time**
AP_ijt	-2.011978 (4)	-1.457708 (4)	-9.014934 (4)a	-9.24750 (4)a
GP_jt	-1.511249 (4)	-1.054364 (4)	-6.952419 (4)a	-7.04114 (4)a
PP_ijt	-1.389683 (4)	-1.239872 (4)	-5.964669 (4)a	-5.98730 (4)a
GARCH_ijt	-1.592908 (4)	-1.437663 (4)	-8.308243 (4)a	-8.36916 (4)a
Indonesia				
Data Series	**without time**	**with time**	**without time**	**with time**
AI_ijt	-2.019012 (4)	-2.792339 (4)	-8.973682 (4)a	-9.00140 (4)a
GI_jt	-1.511249 (4)	-1.054364 (4)	-6.952419 (4)a	-7.04114 (4)a
PI_ijt	-1.966732 (4)	-1.813688 (4)	-5.747611 (4)a	-5.77177 (4)a
GARCH_ijt	-2.66509 (11)	-3.102286 (3)	-5.579170 (4)a	-5.56999 (4)a
Critical values (1% level of significance)				

Notes: Figures in parentheses are the lag order selected based on the SIC where 'a' indicates significance at the 1% significant level.

4.2 JOHANSEN JUSELIUS COINTEGRATION TESTS RESULT

The prerequisite for a set of series to be cointegrated is that they should be integrated in the same order. Thus far, given the power of these unit root tests of I(1) process, now we can proceed to the cointegration test. The cointegration test is designed to test for the presence of common stochastic trends between a set of variable that are individually non-stationary in levels (Hooi, 2007). Thus, in this study, we utilize the Johansen (1988) and Johansen and Juselius (1990) multivariate cointegration process to test for the existence of cointegration relationships, both for trace and maximum eigenvalue statistics. The uniform lag structure of the system is set up through a research process proposed by Vahid and Engle (1993), using the likelihood ratio test with a potential lag length of 1 through 12. The null hypothesis is a system of variables generated from a Gaussian VAR with p_0 lags against the alternative specification of p_1, whereas p_1 is larger than p_0. The test statistic computed is asymptotically distributed as Chi-square with n^2 (p_1-p_0) degrees of freedom. Based on the procedure mentioned above, the optimum lag length of VAR, VECM, and the cointegration for each country are as follows;

Table 2: The Optimum Lag Length for Multivariate Estimation

Systems	Country	Optimum lag length
VAR	Singapore	12
	Malaysia	7
	Thailand	12
	Philippine	9
	Indonesia	6
VECM	Singapore	11
	Malaysia	6
	Thailand	11
	Philippine	8
	Indonesia	5
Cointegration	Singapore	12
	Malaysia	7
	Thailand	12
	Philippine	9
	Indonesia	6

The summary results of cointegration analysis for each county are reported in Tables 3.1 to 3.5. As stated earlier, the number of cointegration vector(s) is determined by two likelihood ratio test, namely maximum eigenvalue and trace eigenvalue statistics. The critical values for each test are from Osterwald-Lenum (1992). Overall, we found similar results for every country under consideration. We conclude that there exists at least one cointegration vector in the cointegration systems. This conclusion applies for Singapore, Malaysia, Thailand, the Philippines and Indonesia.

As an example, for the case of Singapore, the result of trace statistic test obviously demonstrate that the null hypothesis of r=0 against its alternative r>1, is easily rejected at 0.01 and 0.05 significant levels. The computed value of 77.07210 is obviously larger then the 0.05 and 0.01critical values at, 68.52 and 76.07, respectively. Nonetheless, if we test the null hypothesis of r≤1, we definitely fail to reject the hypothesis due to the computed value at 39.36122 is smaller then the 0.05 and 0.01 critical values at, 47.21 and 54.46, respectively. Therefore, based on the trace statistic test results, we conclude that there exists a single cointegrating vector in the model. Consistent with that, we also found the maximum eigenvalue test suggests a similar result.

Moreover, based on the trace and maximum eigenvalue tests, the results reveal the null hypothesis of r=0 against its alternative r>1 is rejected at 0.01 and 0.05 significant levels. Using the EViews v6 software as our estimating engine, the computed value 37.71088 is obviously larger then the 0.05 and 0.01 critical values at, 33.46 and 38.77, respectively. Nonetheless, if we test the null hypothesis of r≤1, we definitely fail to reject the hypothesis due to the computed value at 15.98193 is smaller then the 0.05 and 0.01 critical value at, 27.07 and 32.24, respectively.

Overall, summarizing the case of Singapore, there is at least one cointegrating vector in the system. Based on this conclusion, this study furthermore suggests that the economic growth and its macroeconomic determinants exhibit a long-run relationship in the Singapore cointegrating system. This is means the series in that system move together and can not move far from each other. The same conclusions can also be applied for the cases of Malaysia, Thailand, the Philippines and Indonesia.

Table 3.1 : The cointegration results for Singapore

Hypothesis		λ Trace	5% critical value	1% critical value	λ Max	5% critical value	1% critical value
H0	H1						
r=0	r>0	77.07210$^{*(**)}$	68.52	76.07	37.71088*	33.46	38.77
r≤1	r>1	39.36122	47.21	54.46	15.98193	27.07	32.24
r≤2	r>2	23.37929	29.68	35.65	13.10595	20.97	25.52
r≤3	r>3	10.27334	15.41	20.04	8.859741	14.07	18.63
r≤4	r>4	1.413595	3.76	6.65	1.413595	3.76	6.65

Note that, the notation 'r' denotes the number of cointegrating vectors. The superscript (**) indicates statistically significant at 5% and (*) at 1%. The critical values for the Johansen Juselius test were obtained from (Osterwald-Lenum, 1992).

Table 3.2 : The cointegration results for Malaysia

Hypothesis		λ Trace	5% critical value	1% critical value	λ Max	5% critical value	1% critical value
H0	H1						
r=0	r>0	88.36177$^{*(**)}$	68.52	76.07	41.53120$^{*(**)}$	33.46	38.77
r≤1	r>1	46.83056	47.21	54.46	18.54683	27.07	32.24
r≤2	r>2	28.28373	29.68	35.65	13.41724	20.97	25.52
r≤3	r>3	14.86649	15.41	20.04	11.49842	14.07	18.63
r≤4	r>4	3.368069	3.76	6.65	3.368069	3.76	6.65

Note that, the notation 'r' denotes the number of cointegrating vectors. The superscript (**) indicates statistically significant at 5% and (*) at 1%. The critical values for the Johansen Juselius test were obtained from (Osterwald-Lenum, 1992).

Table 3.3 : The cointegration results for Thailand

Hypothesis		λ Trace	5% critical value	1% critical value	λ Max	5% critical value	1% critical value
H0	H1						
r=0	r>0	101.5702$^{*(**)}$	68.52	76.07	58.44194$^{*(**)}$	33.46	38.77
r≤1	r>1	43.12830	47.21	54.46	16.92207	27.07	32.24
r≤2	r>2	26.20623	29.68	35.65	14.05542	20.97	25.52
r≤3	r>3	12.15081	15.41	20.04	9.563480	14.07	18.63
r≤4	r>4	2.587330	3.76	6.65	2.587330	3.76	6.65

Note that, the notation 'r' denotes the number of cointegrating vectors. The superscript (**) indicates statistically significant at 5% and (*) at 1%. The critical values for the Johansen Juselius test were obtained from (Osterwald-Lenum, 1992).

Table 3.4 : The cointegration results for the Philippines

Hypothesis							
H0	H1	λ Trace	5% critical value	1% critical value	λ Max	5% critical value	1% critical value
r=0	r>0	86.41198$^{*(**)}$	68.52	76.07	40.81620$^{*(**)}$	33.46	38.77
r≤1	r>1	45.59578	47.21	54.46	17.46097	27.07	32.24
r≤2	r>2	28.13480	29.68	35.65	14.86239	20.97	25.52
r≤3	r>3	13.27241	15.41	20.04	9.544347	14.07	18.63
r≤3	r>3	3.728066	3.76	6.65	3.728066	3.76	6.65

Note that, the notation 'r' denotes the number of cointegrating vectors. The superscript (**) indicates statistically significant at 5% and (*) at 1%. The critical values for the Johansen Juselius test were obtained from (Osterwald-Lenum, 1992).

Table 3.5 : The cointegration results for Indonesia

Hypothesis							
H0	H1	λ Trace	5% critical value	1% critical value	λ Max	5% critical value	1% critical value
r=0	r>0	86.93364$^{*(**)}$	68.52	76.07	47.10336$^{*(**)}$	33.46	38.77
r≤1	r>1	39.83028	47.21	54.46	19.36090	27.07	32.24
r≤2	r>2	20.46938	29.68	35.65	12.46644	20.97	25.52
r≤3	r>3	8.002940	15.41	20.04	5.369981	14.07	18.63
r≤3	r>3	2.632959	3.76	6.65	2.632959	3.76	6.65

Note that, the notation 'r' denotes the number of cointegrating vectors. The superscript (**) indicates statistically significant at 5% and (*) at 1%. The critical values for the Johansen Juselius test were obtained from (Osterwald-Lenum, 1992).

4.3 ERROR CORRECTION MODELS RESULT

In this part, since the cointegration tests in the earlier section identify for the one long-run relationship for each of the export equation, the error correction models were estimated to observe the short-run relationship in the models[6]. In order to find a reasonable equation for exports models, we performed many estimation experiments. In this stage, there will be two parts of the estimation procedure (Baak, 2008). Firstly, the estimated regression equation includes each explanatory variable up to 12 lags. Then, each lagged variable which was not found to be insignificant will be omitted from the systems. The full results for this stage are reported in Table 4.

In table 4, the results suggest for the long-run equilibrium relationship among the variables in each export function. These results are further supported with the negative sign of the each of error terms coefficient (ECT_{ijt-1}) in the exports function. Additionally, all systems passed the diagnostic tests. Here, we perform the Durbin-Watson tests, Breusch-Godfrey tests and Breusch-Pagan-Godfrey tests as diagnostic tests. Moreover, some of the estimated coefficients of the explanatory variables were consistent with the previous studies, namely, Baak et al. (2007), Chit et al. (2008), Arize et al. (2008) and Baak (2009). Nevertheless, the results overall contradict our expectations in just a few cases. Firstly, the effects of the GDP of the importing country (in our study this is the United States), are estimated to be both negative and positive in the systems. But, for Thailand the relationship between exports and GDP are positive. Therefore, the overall effects of the relationship are positive and ambiguous. Secondly, the result suggests a positive relationship in the short-run between exports and the bilateral exchange rate, for Singapore and Malaysia. This result denotes that, when depreciation of the exporting country's currency (depreciations of the domestic currency i.e; Ringgit Malaysia (RM) for the Malaysia case) usually leads to an increase in exports (from the United States). However, this finding does not apply for Indonesia, Thailand and Philippines, where the results are mixed and lead to sign ambiguity.

Third, the short-run effects of the exchange rate volatility are more complicated. There are positive effects in the exports of Indonesia to the United States. Besides, the results further suggest for the negative relationship between exports and exchange rate volatility, from Philippines and Singapore to the United States. The results are found to be mixed in the Malaysia and Thailand systems. Therefore, as a conclusion, the effects of the exchange rate volatility to Malaysia and Thailand are ambiguous, while the same relationship for Indonesia

[6] The error correction terms were estimated by the cointegrating equation systems. The summary results were reported in the Table 4.

and the Philippines/Singapore are positive and negative, respectively. Finally, the table also shows significant effects from the crisis dummy to exports. Therefore, to take into account the crisis dummy in the systems is vital in order to capture the structural break that occurred during the 1997/1998 Asian Financial crisis.

5 CONCLUSION

This paper offers some new results for the exchange rate volatility from ASEAN countries to the United States over the monthly period from January, 1990 to December, 2010. In order to capture for the short and long run relationship between the variables under estimation, this study performed the Johansen Juselius (1990) tests and error correction model in order to capture for the short- and long-run relationship between the variables in the systems. In general, the real bilateral exchange rate volatility has a significant impact on exports at least for all the countries considered in our sample, and the impact overall is negative except for Indonesia.

Table 4: Error correction mechanisms dependent exports
Sample: January, 1990- December, 2010

Variables	ASEAN Countries				
	Singapore	Malaysia	Thailand	Philippine	Indonesia
Constant	-0.0009 (-1.58)	0.006 (1.66)c	0.018 (4.674)a	0.004 (0.99)	0.012 (2.29)b
ECT_{ijt-1}	-0.2296 {-5.00}a	-0.018{-1.78}c	-0.0254{-7.42}a	-0.028 {-5.70}a	-0.013 {-1.57}c
ΔA_{ijt-1}	-0.681 (-11.33)a	-0.279 (-4.41)a	-0.599 (-9.56)a	-0.375 (-6.18)a	-0.444 (-7.46)a
ΔA_{ijt-2}	-0.262 (-3.80)a	-0.088 (-1.59)	-0.557 (-7.66)a	-0.324 (-5.34)a	-0.203 (-3.20)a
ΔA_{ijt-3}	0.104 (1.57)	-	-0.438 (-6.19)a	-0.177 (-3.21)a	-0.175 (-2.926)a
ΔA_{ijt-4}	0.161 (2.73)a	0.019 (0.34)	-0.508 (-7.97)a	-0.302 (-4.94)a	-0.194 (-3.14)a
ΔA_{ijt-5}	-	-	-0.463 (-7.43)a	-0.271 (-4.41)a	-0.222 (-3.521)a
ΔA_{ijt-6}	-	-0.138 (-2.32)b	-0.526 (-8.95)a	-0.214 (-3.74)a	-0.226 (-3.697)a
ΔA_{ijt-7}	-0.199 (-3.55)a	-0.144 (-2.40)b	-0.581 (-9.65)a	-0.156 (-2.82)a	-0.109 (-1.705)c
ΔA_{ijt-8}	-	-	-0.584 (-9.29)a	-0.268 (-4.31)a	-0.087 (-1.47)
ΔA_{ijt-9}	0.102 (1.77)c	-	-0.482 (-7.22)a	-0.28 (-4.65)a	-
ΔA_{ijt-10}	0.100 (1.72)c	-	-0.404 (-5.98)a	-0.187 (-3.39)a	-0.093 (-1.761)c
ΔA_{ijt-11}	-	0.129 (2.17)b	-0.248 (-4.23)a	-	-
ΔA_{ijt-12}	0.144 (2.61)a	0.257 (4.27)a	-	0.206 (3.48)a	0.242 (4.44)a
ΔG_{jt-1}	0.009 (2.57)b	0.008 (2.86)a	0.008 (2.86)a	0.016 (4.40)a	0.015 (3.48)a
ΔG_{jt-2}	0.020 (4.69)a	0.016 (4.62)a	0.019 (6.22)a	0.012 (3.316)a	0.014 (3.59)a
ΔG_{jt-3}	0.012 (3.21)a	0.002 (0.69)	0.011 (3.08)a	-	-
ΔG_{jt-4}	-	-	0.007 (2.16)b	-	-
ΔG_{jt-5}	-0.013 (-3.38)a	-	0.007 (2.44)b	0.003 (1.27)	0.009 (2.29)b
ΔG_{jt-6}	-0.021 (-5.34)a	-0.005 (2.12)b	-	-	-0.007 (-1.93)c
ΔG_{jt-7}	-	-	0.012 (4.21)a	-	-
ΔG_{jt-8}	-	-0.015 (-4.92)a	-	-0.015 (-4.70)a	-0.0089 (-2.19)b
ΔG_{jt-9}	0.014 (3.64)a	-0.005 (-1.68)c	0.013 (4.49)a	0.005 (1.67)c	-0.004 (-1.08)
ΔG_{jt-10}	0.007 (1.91)c	-	0.009 (3.01)a	-	-
ΔG_{jt-11}	-	-	-	-	-
ΔG_{jt-12}	-	-	-	0.009 (2.93)a	0.008 (2.09)b
ΔP_{jt-1}	-	-	-	-	0.155 (1.79)b
ΔP_{jt-2}	0.962 (2.60)b	-	0.013 (0.08)	0.262 (1.36)	-
ΔP_{jt-3}	-	0.246 (1.13)	-	-	-
ΔP_{jt-4}	0.753 (2.06)b	-	-	-	-
ΔP_{jt-5}	-	-	-	0.544 (2.48)b	-0.316 (-2.27)b
ΔP_{jt-6}	0.544 (1.45)	0.506 (2.32)b	-0.01 (-0.07)	-0.256 (-1.29)	-
ΔP_{jt-7}	0.839 (2.18)b	-	-	-	-
ΔP_{jt-8}	-	-	-	-	-
ΔP_{jt-9}	-	0.49 (2.418)b	-	-	-
ΔP_{jt-10}	0.429 (1.13)	-	0.214 (1.41)	-	-
ΔP_{jt-11}	-	-	-	0.240 (1.23)	-0.174 (-2.16)b
ΔP_{jt-12}	-	-	-	-	-
$\Delta \sigma_{ijt-1}$	-	-	-	-	0.148 (1.17)
$\Delta \sigma_{ijt-2}$	-	-	0.620 (2.52)b	-	0.007 (0.063)
$\Delta \sigma_{ijt-3}$	-	0.986 (1.43)	0.291 (0.89)	-	-
$\Delta \sigma_{ijt-4}$	-	-	-	-2.02 (-3.13)a	0.500 (2.499)b
$\Delta \sigma_{ijt-5}$	-	-1.215 (-1.65)c	-	-	-
$\Delta \sigma_{ijt-6}$	-2.09 (-1.15)	-	-	-0.818 (-1.39)	-
$\Delta \sigma_{ijt-7}$	-4.23 (-2.26)b	-1.135 (-1.460)	0.103 (0.45)	-1.189 (-2.08)b	-
$\Delta \sigma_{ijt-8}$	-	-0.737 (-0.97)	-	-	-
$\Delta \sigma_{ijt-9}$	-	-	-	-	-
$\Delta \sigma_{ijt-10}$	-	-	-0.086 (-0.24)	-1.672 (-2.67)a	-
$\Delta \sigma_{ijt-11}$	-	-	-	-	-
$\Delta \sigma_{ijt-12}$	-	1.118 (1.93)c	-	-	-
ΔCD_{ijt-1}	-	-0.028 (-1.65)	-	-	-
ΔCD_{ijt-2}	-	0.111 (2.62)a	0.134 (3.16)a	0.105 (2.37)b	0.153 (2.69)a
ΔCD_{ijt-3}	-	-0.066 (-1.54)	-	-0.093 (-2.09)b	-
ΔCD_{ijt-4}	-0.126 (-2.24)b	-0.104 (-2.43)b	-0.037 (-0.91)	-	-0.072 (-1.28)
ΔCD_{ijt-5}	-0.094 (-1.67)c	-0.058 (-1.309)	-	-0.034 (-0.74)	-
ΔCD_{ijt-6}	-0.082 (-1.45)	-	-	-	-
ΔCD_{ijt-7}	-0.122 (-2.12)b	-	-0.006 (-0.14)	-0.094 (-1.98)b	-0.102 (-1.65)c
ΔCD_{ijt-8}	-	-	-	-0.059 (-1.258)	-
ΔCD_{ijt-9}	-	-	-	-	-
ΔCD_{ijt-10}	-	-	-	-	-

Table 4: Regression results for Error Correction Models (Continued)

Variables	Country				
	Singapore	Malaysia	Thailand	Philippine	Indonesia
ΔCD_{ijt-11}	-	-	0.075 (1.44)	-	-0.094 (-1.46)
ΔCD_{ijt-12}	-	0.106 (2.25)b	0.004 (0.10)	-	-
DW	2.0213	2.0680	1.9032	2.0800	2.0822
B-G	**F=1.5[lag4/0.34]**	**F=1.9[lag12/0.35]**	**F=1.2[lag2/0.31]**	**F=1.1[lag2/0.22]**	**F=1.5[lag2/0.25]**
B-P-G	**F=2.25[0.1887]**	**F=2.3[0.2432]**	**F=1.69[0.2938]**	**F=2.03[0.2180]**	**F=1.12[0.3180]**
R^2 **Ads. R^2** **F-stat** **Prob.(F-Stat)**	**0.6343** **0.5868** **13.3624** **(0.0000)**	**0.5384** **0.4759** **8.6229** **(0.0000)**	**0.6816** **0.6296** **13.1045** **(0.0000)**	**0.6306** **0.5744** **11.2332** **(0.0000)**	**0.5746** **0.5171** **9.9880** **(0.0000)**
5% Critical values					

Note that, figure in parentheses are the absolute t-statistic.

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