Asymmetric Co-integration between Exchange Rate and Trade Balance in Thailand

Saad Buba¹, Salim Mohammed Ibrahim Al-Jadi² Garba Mohammed Guza³

¹Adamawa State Polytechnic, Yola, Nigeria.
²Faculty of Economics, Sabha University, Libya.
³Mai Idriss Alooma Polytechnic, Geidam, Nigeria.

This paper empirically examines the long-run exchange rate pass-through into trade balance in Thailand. The study incorporates political stability in the short run model to ascertain its effect on the trade balance. Asymmetric co-integrating adjustment method proposed by Enders and Siklos (2001) is employed for the study. The empirical findings revealed that there exists an asymmetric cointegration relationship between exchange rate and trade balance as well as exchange rate and imports & exports volumes after conducting a momentum-threshold autoregressive (M-TAR) and threshold autoregressive (TAR) tests respectively. The results of the short run effects showed that political stability has no meaningful effect on trade balance of Thailand. The findings have further shown that changes in real exchange rate have contributed to the presence of trade balance deficit in Thailand during the period under study; which is most likely to be as a result of massive imports of crude oil between the late 1990’s and through to 2010.

Keywords: Asymmetric adjustment; Exchange rate; Trade balance; Thailand; Import and Export

INTRODUCTION

In Thailand, just like in other developing countries, the trade deficit is one of its major hiccups. Import Prices relate to the rate of variation in the prices of goods and services bought by people of the country from foreign sellers and delivered by same foreign suppliers. A question may be asked, do import prices greatly affected by exchange rates. In January 2015 import prices in Thailand were 91.65 index points, one month later, it declined to 91.41 Index Points. Import prices in Thailand, from 2000 to 2015, averaged at 88.46 Index Points, getting an all-time, high to 103.35 Index Points in January 2014 and the lowest at 67.79 Index Points in January 2000 (Bank of Thailand, 2015 report). The response of international prices to changes in exchange rate remains a central issue, the exchange rates pass-through to import prices dictates the degree of change to the current account and international transmission of inflation, (Cook, 2014).

In 1999, after the economic and financial crisis in 1997, effort has been made by the Royal Thai Government to stabilize the economy of Thailand. In 2003, the economy was set back on track; the exchange rate stability has been achieved, (Trade Policy Review report, 2003). Between 2000 and 2009, the average exchange rate of Baht to US dollar depreciated and the trade deficit emerge due to the absence of effective fiscal policy. The economist report 2014 revealed that Thailand would remain the slowest growing economy in South-East Asia, probably till 2017. Melitz (2003), in his model of heterogeneous firms, predicts that a change in trade conditions causes a change in the variety of goods that are traded.

*Corresponding author: Saad Buba, Adamawa State Polytechnic, Yola, Nigeria. E-mail: saadbubadamare12@gmail.com
This can be explained by the fact that when trade costs decreased, firms that were unable to export before the decrease will now do so. But the new firms’ productivity will be lower or will have higher marginal cost than the firms already exporting before the decrease. In addition, the new firms with lower productivity also have a pricing behavior that will determine how aggregate pass-through to a country’s import prices changes resulting from a decrease in trade costs.

In recent years, Thailand has recorded a fall in its trade balance from USD 36.6 billion in 2009 to USD 6.8 billion in 2013, making a short of USD 29.8 billion (Focus Economics, 2015). A question may be asked; is it the depreciation of Baht to US dollar that contributed to this trade balance deficit faced by Thailand? This study will examine the exchange rate pass-through into trade balance in Thailand. To the best of our knowledge, there is no study on Thailand that examined exchange rate pass-through into trade balance using asymmetric cointegration with threshold autoregressive (TAR) and momentum threshold autoregressive (M-TAR). It is imperative, therefore, to study this pass-through of exchange rates into trade balance in Thailand using asymmetric cointegration approach proposed by Enders and Siklos (2001).

Previous studies have used cointegration analysis to examine the relationship between exchange rates and trade balance such as the Engle-Granger (1987), two-step cointegration analysis and Johansen Juselius (1999), to examine the long run relationship between exchange rate and trade balance, see (Hsing 2009), Wong and Tang (2007), Bagchi et.al (2004), Khan (2005). The Engle-Granger (1987) and Johansen Juselius (1999), assumed symmetric cointegration; they show the changing coefficient values related either negatively or positively in the symmetry error. Studies dealing with symmetry relationship do not consider the account of asymmetric adjustment that occurs in the endogenous variables, (Duasa, 2009). Precisely, Cook (2006) argued that Engel and Granger (1987) and Johansen (1988) were unable to identify the cointegration between stock prices and industrial production for a long span of U.S data. However, Narayan (2007) has employed asymmetric approach using TAR and M-TAR for nominal exchange rates and price levels for Canada, the UK, Japan, Germany, Italy and France (G6) vis-à-vis the US dollar. Shen et.al (2007) has employed the asymmetric cointegration test proposed by Enders and Siklos (2001) to examine the long run asymmetric equilibrium relationships between the Chinese Shanghai and Shenzhen stock markets. The asymmetric relationship entails differentiating positive and negative effects of the error obtained from the cointegration regression, (Enders and Siklos, 2001). Asymmetry has been a significant matter in macroeconomics analysis with findings giving proof of the existence of asymmetric changes of most of the macroeconomic indicators.

This study is theoretically, based on the standard theory of international trade by Heckscher-Ohlin (1990’s), which states that changes in real exchange rate yield impacts on both the value and volume of trade; the higher the real exchange rate for the home country, the more the trade surplus the country obtains and vice-versa. The major different between this study and Duasa (2009) besides the choice of country of study is the ability of the current study to incorporate political stability in the short-run model and the idea employed by the study to conduct robust checks using fully modified ordinary least squares (FMOLS) and the dynamic ordinary least squares (DOLS).

RELATED LITERATURE

There is various literature concerning trade balance, Kayhan et al. (2013) have examined the causality between the trade deficit and government expenditure in the Turkish economy; they employed Toda-Yamamoto causality (bootstrap process based) and frequency domain analysis method, their findings revealed a bi-directional causality between the trade deficit and government expenditure. This is evident that each method applied in their study is robust; government expenditure affects trade deficit in the shorter period, whereas trade deficit affects government expenditure even in the long run. They conclude based on theoretical and empirical interpretation that reduction in government expenditure will help reduce the trade deficit. A more recent study was conducted by Prakash and Maiti (2016) who investigated the devaluation effectiveness in improving trade balance in Fiji. They used econometric models to empirically examine devaluation impact; they found that real exchange rate has a strong relationship with trade balance in the long run. Appreciation of foreign currency is responsible for the rising trade deficit. This is due to Fiji’s over-dependent on imports and its failure to cure import bills by raising domestic export. Another recent study by Chiu and Sun (2016) examines the relationship between trade balance, savings and exchange rates in a panel of 76 countries. The study uses panel smooth transition regression model with instrumental variables to address the potential non-linear effects of the savings rate. They found that countries with savings rate above 14.8 percent threshold can induce their trade balance by increasing the savings rate.

Aliyu and Tijjani (2015) examined the long-run exchange rates pass-through into trade balance in Nigeria by means of threshold cointegration and asymmetric error correction modelling. Niemenien (2015) found that the introduction of a common currency in the EU area did not increase the elasticity of net capital flows to per capita income; hence the understanding of imbalance in the Euro area increased. Interestingly, countries with the most sophisticated financial markets have had positive intra balances and negative extra balances. Whalley and Wang (2011) found that the effect of Renminbi revaluation on
surplus is proportionally larger than on trade flows; changes in trade flows can be substantial. China’s processing trade was incorporated into their model, they found that the impact of revaluation on the surpluses are elasticity dependent; huge substitution elasticities between domestic and foreign production in demand in China yield larger effects on trade, its flow and surpluses. Devaluation of US dollar deteriorates her bilateral trade balance with 13 trading partners, especially for China, and the U.S. trade imbalance can be eliminated by way of dollar devaluation in the long run, (Chiu et al. 2010). Contrary to many empirical studies but consistent with the theoretical expectation regarding the effect of the contractionary policy, Ivrendi and Guloglu (2010) found that monetary-elasticities of the rate of real exchange for exports and imports, respectively, while the inability to compel into consideration, the currently acknowledged evidence of the existence of asymmetric changes of macroeconomic variables, may lead to inappropriate interpretations (Duasa 2009). This is because according to Balke and Fomby, (1997), the movement towards the long-run steady state will not be essentially constant, Hence the need for this study. The rest of the work of this paper is structured as follows: the theoretical framework is presented in the next section, that is section 3, empirical methods are presented in section 4; the highlights of the empirical findings and analysis are presented in section 5. Section 6, the final section concludes and draws policy recommendation from the major findings.

Theoretical Framework

In this section, we provide an overview of the theoretical relationship between exchange rate and trade volumes (import & export volumes). Rose and Yellen (1989) and Rose (1990) presented a model which specified in a reduced form, the performance of trade balance in a given economy as equation (1) and (2) below. The model is presented based on the real exchange rate and real incomes generated within and outside the economy.

\[ X_t = \left( \frac{\mu}{\rho \varepsilon} \right) \cdot (Y_t) \varepsilon \]  
\[ M_t = \left( \frac{\rho^*}{\rho} \right) \cdot (Y_t) \varepsilon^* \]  

According to the models, \( X \) denotes exports, \( M \) is imports, \( E \) is the minimum exchange rate (as in for example Duasa (2009)) and \( P, P^* \) and \( Y, Y^* \) indicate the domestic and foreign levels of prices and incomes; \( \eta \) and \( \gamma \) denote the elasticities of the rate of real exchange for exports and imports, respectively, while \( \varepsilon \) stands for export’s elasticity of income, \( z \) is the income elasticity for imports. We rewrite equations (1) and (2) by taking their natural log:

\[ \ln X_t = \eta [ \ln P_t - \ln P^*_t - \ln E_t ] + \varepsilon \ln Y_t^* \]  
\[ \ln M_t = \gamma [ \ln P^*_t - \ln E_t - \ln P_t ] + z \ln Y_t \]  

Where \( \ln e_t = [ \ln P^*_t + \ln E_t - \ln P_t ] \) designates natural log of real exchange rate. Under normal condition, balance of trade \( (TB) \) is the ratio of exports to imports or otherwise and its equation could be written as in Duasa (2009).

\[ \ln TB_t = z \ln Y_t + \varepsilon \ln Y_t^* + \theta \ln e_t \]  

Where \( \theta = - (\eta + \gamma) \). The amount of \( \ln e_t \) shows the condition of Marshal-Lerner (M-L), whether achieved or not. The M-L condition occur when exchange rate devaluation will improve balance of trade as the sum of long-term export and import demand elasticities less than one, (Davidson, 2009). In order to achieve the M-L condition, \( \eta \) and \( \gamma \) are expected to be non-positive then \( \varepsilon \) and \( z \) are expected to be non-negative hence, M-L holds at any time \( \theta \) is non-negative showing that rate of exchange is higher.

Empirical Approach

Enders and Siklos (2001), have expanded Engle-Granger two-step cointegration tests, to integrate an asymmetric error correction term, assumed two series\( \{ y_t, x_t \} \). The ordinary least squares (OLS) is employed in the first step to examine the symmetric correlation among the \( y_t \) and \( x_t \) in the long-run; thus:

\[ y_t = \hat{\theta}_0 + \hat{\theta}_1 x_t + \varepsilon_t \]  

Where \( \hat{\theta}_0 \) and \( \hat{\theta}_1 \) designate the parameters to be estimated and \( \varepsilon_t \) is an error term. The conceivable cointegration among \( y_t \) and \( x_t \) is then estimated through the order of incorporation of the residuals \( \varepsilon_t \) in equation (6) by means of a Dickey-Fuller test as follows:

\[ \Delta \hat{\varepsilon}_t = \rho \hat{\varepsilon}_{t-1} + \nu_t \]  

By way of suitable degree of augmentation put in place through the insertion of lagged values of the endogenous variable, \( (H_0: \rho = 0) \) as null hypothesis of no cointegration can be properly examined through the assessment of the \( t \)-ratio of the changing parameter \( \rho \) and precisely produced unusual critical values. Though, according to Enders and Siklos (2001), the misspecification of Engle-Granger method occurs when the estimated time series have a fundamental asymmetric correlation. Hence, stationary testing of the error term is done by integrating the asymmetric changes. A model was proposed by Enders and Siklos (2001) in the second step as follows:

\[ \Delta \varepsilon_t = \lambda_t \rho \varepsilon_{t-1} + (1 - \lambda_t) \rho \varepsilon_{t-1} + \sum_{i=1}^{\infty} \beta \varepsilon_{t-1} + \varphi_t \]
Where $\rho_1$, $\rho_2$ and $\rho_i$ are coefficients; $\phi_i$ is a white-noise disturbance; $k$ is the number of lags; and $I_t$ is an indicator function such that:

$$I_t = \begin{cases} 1 & \text{if } \varepsilon_{t-1} \geq 0 \\ 0 & \text{if } \varepsilon_{t-1} < 0 \end{cases}$$

(9)

Tong (1983, 1990) revealed that the OLS estimates of $\rho_1$ and $\rho_2$ contained an asymptotic multivariate normal distribution. Equations 6, 8 and 9 are cointegration model and are called the Enders and Siklos (2001)’s threshold autoregressive (TAR) model. It remains imperative to know that the indicator function $I_t$ is influenced by the rate of $\varepsilon_{t-1}$ in equation (9).

An alternative threshold has been suggested by Enders and Siklos (2001) and Enders and Granger (1998). This suggested threshold depends on the adjustment in $\varepsilon_{t-1}$ in the preceding period and consequently, the new indicator $M_t$ is emerged as:

$$M_t = \begin{cases} 1 & \text{if } \Delta \varepsilon_{t-1} \geq 0 \\ 0 & \text{if } \Delta \varepsilon_{t-1} < 0 \end{cases}$$

(10)

The model that contains equations 6, 8 and 10, is devoted to the momentum threshold autoregressive (M-TAR) model. The threshold ($\delta$) value is set to 0 in equations 9 and 10. When the threshold ($\delta$) is not known; the use of grid process is suggested by Enders and Siklos (2001), in order to come up with a consistent and dependable estimate of the threshold. To be precise, we bear in mind that TAR model, the residual series $\{\hat{\varepsilon}_t\}$ is orderly and arranged as $\{\hat{\varepsilon}_1 < \hat{\varepsilon}_2 < \ldots < \hat{\varepsilon}_T\}$. After thrown away 15 percent of the largest and smallest $\{\hat{\varepsilon}_t\}$, the dominant observations which is 70 percent of the array are formally reflected as thresholds in equations (8) and (9) as every single of the observation could be the conceivable threshold. The estimated threshold considered yielding the lowest residual sum of squares stood to be believed as the right estimate of the threshold. A related method can be applied for M-TAR model. With 70 percent of observations of the sequence $\{\hat{\varepsilon}_1 < \hat{\varepsilon}_2 < \ldots < \hat{\varepsilon}_T\}$ measured as values of the threshold for equation (10), possibly, the value may offer least sum of residual squares emanating from the estimation of equations (8) and (10), is then described as the consistent threshold.

To check for asymmetric cointegration; firstly, we determine whether or not $y_t$ and $x_t$ are cointegrated in the TAR and M-TAR models. We carry out this test by means of $F$ test meant to investigate the null hypothesis of no cointegration, $H_0: \rho_1 = \rho_2 = 0$. Enders and Siklos (2001) denoted $F$ as the $F$ statistics and it has a non-standard distribution. Secondly, the null hypothesis $H_0: \rho_1 = \rho_2$ can be tested using the standard $F$-statistics, however this can be done in the existence of asymmetric cointegration. The evidence supporting the asymmetric adjustment of the error correction term is specified when both $H_0: \rho_1 = \rho_2 = 0$ and $H_0: \rho_1 = \rho_2$ are rejected. Possible adjustment is $\rho_1$ if $y_{t-1}$ is above its long run equilibrium value ($= y_0 + y_1 x_{t-1}$), but $\rho_2$ if $y_{t-1}$ is below the equilibrium.

The error-correction model of the asymmetric correspondingly exists for $y_t$ and $x_t$ once they are made in an asymmetric cointegration relationship. Thus:

$$\Delta y_t = \alpha_0 + \theta_1 M_t \varepsilon_{t-1} + \theta_2 (1 - M_t) \varepsilon_{t-1} + \sum_{i=1}^k \alpha_i \Delta y_{t-i} + \sum_{i=1}^k \alpha_{2i} \Delta x_{t-i} + \vartheta_{1t}$$

(11)

And

$$\Delta x_t = \beta_0 + \theta_1 M_t \varepsilon_{t-1} + \theta_2 (1 - M_t) \varepsilon_{t-1} + \sum_{i=1}^k \beta_i \Delta y_{t-i} + \sum_{i=1}^k \beta_{2i} \Delta x_{t-i} + \vartheta_{2t}$$

(12)

Where $\theta_{11}$ and $\theta_{12}$ signify the adjustment coefficients’ speed of $\Delta y_t$ if $y_{t-1}$ is above and below its long run steady state, respectively. Likewise, $\theta_{21}$ and $\theta_{22}$ denote the adjustment coefficients’ speed of $\Delta x_t$ of the two systems, respectively, $\alpha_0$ and $\beta_0$ are the constant terms, $\alpha_{11}$, $\alpha_{21}$, $\beta_{12}$ and $\beta_{22}$ are coefficients of lagged change terms, and $\vartheta_{1t}$ and $\vartheta_{2t}$ are white-noise disturbances.

In addition, we apply the Granger causality test in order to observe the lead-lag correlation among $y_t$ and $x_t$. The null hypothesis; $x_t$ does not lead $y_t$ is $H_0: \alpha_{21} = 0, i = 1, \ldots, k$ and the null hypothesis that $y_t$ does not lead $x_t$ is $H_0: \beta_{12} = 0, i = 1, \ldots, k$.

**EMPIRICAL RESULTS AND ANALYSIS**

**Data description and unit root test**

The observed series of trade variables in this study are all together, collected from the Bank of Thailand (BOT)’s data and statistics database. A monthly data of fifteen years, from M1: 2000 to M12: 2014; see Chen, et al. (2005) and Ibrahim (2011) on the use of the data sample size. The measure of the exchange rate is the exchange rate of commercial banks in Bangkok, with the base year of 2000. The imports and exports volumes are both in indices with 2000 as the base year, and finally, the balance of trade is basically the difference between export volume and import volume, (Luo, Lv & Zhang 2010). Imports and exports values are in log form, except for the trade balance. Political stability data were taken from World Bank’s world governance indicator (WGI); an annual data resampled to monthly using Eviews.

The first test conducted in the unit root test; breakpoint unit root, in particular, using a modified Dickey-Fuller test with Eviews 9 software. It carries out a test which allows for a structural break in the process. The latest Eviews 9 supports the computation of modified Dickey-Fuller tests which allow for levels and trends that differ across single break date. The framework follows the work of Perron (1989), Perron and Vogelsang (1992), Vogelsang and Perron (1998), Banerjee, et al. (1992). The results of the unit root tests are displayed in Tables 1 and 2 below.
We present the following functions which are considered significant. The effect of its break dummy variable is positive (1.466) and significant at five percent. Lastly, the effect of the break dummy variable on the variable is also positive but insignificant. The break date for the trade balance is 2009, second month and the effect of its break dummy variable is positive (0.1748) and significant at one percent. The break date for the import volume has its break date in 2012, third month; the effect of its break dummy variable on it is also positive but insignificant. The export volume is 2008, seventh month; the effect of its break dummy variable is negative but not significant. The import volume has its break date in 2012, third month; the effect of its break dummy variable is positive (0.331) and significant at one percent. The break date for the real exchange rate is 2001, t = 15.

Table 1: Breakpoint Unit Root Tests

<table>
<thead>
<tr>
<th>Variable</th>
<th>Minimize Dickey-Fuller test statistics</th>
<th></th>
</tr>
</thead>
<tbody>
<tr>
<td>Tb</td>
<td>-3.9798</td>
<td>-6.6659***</td>
</tr>
<tr>
<td>X</td>
<td>-3.6053</td>
<td>-5.2154**</td>
</tr>
<tr>
<td>M</td>
<td>-3.1415</td>
<td>-5.9743***</td>
</tr>
<tr>
<td>rer</td>
<td>-3.9745</td>
<td>-10.128***</td>
</tr>
</tbody>
</table>

Note: *, **, and *** indicate level of significance at 10%, 5% and 1% level, respectively.

Table 2: Breakpoint Unit Root Test

<table>
<thead>
<tr>
<th>Variable</th>
<th>Minimize Intercept break test statistics</th>
<th></th>
</tr>
</thead>
<tbody>
<tr>
<td>Tb</td>
<td>-4.6409*</td>
<td>-16.600***</td>
</tr>
<tr>
<td>X</td>
<td>-2.9157</td>
<td>-4.7754**</td>
</tr>
<tr>
<td>m</td>
<td>-4.0771</td>
<td>-22.1629***</td>
</tr>
<tr>
<td>rer</td>
<td>-4.1997</td>
<td>-9.5069***</td>
</tr>
</tbody>
</table>

Note: *, **, and *** indicate level of significance at 10%, 5% and 1% level, respectively.

Our test results for two structural breaks (minimize Dickey-Fuller t-statistics and minimize intercept break t-statistics) showed statistics coefficient with p-values less than 0.01 all at first difference except for trade balance under minimize intercept break t-statistics. This has led us to reject the null of a unit root. The break dates and the effect sign of break dummy variable on each of our variables and their significance levels were also provided.

For real exchange rate, the break date is 2001, the fifth month is shown by both Dickey-Fuller t-statistics and autoregressive coefficient graphs; the effect of break dummy variable on the exchange rate is negative but not significant. The import volume has its break date in 2012, third month; the effect of its break dummy on it was positive (0.331) and significant at one percent. The break date for export volume is 2008, seventh month; the effect of its break dummy variable is positive (0.1748) and significant at five percent. The trade balance has a break date in 2009, second month and the effect of the break dummy variable on it is also positive but insignificant. Lastly, political stability has a break date in 2007, sixth month; the effect of its break dummy variable is positive (1.466) and significant at ten percent. See Appendix 1 and 2 for break dates and break dummy variables effects.

The long run relations

We present the following functions which are considered suitable for long run correlation between exchange rate and trade balance, export volume & import volume, as in the work of Duasa (2009):

Model 1: \[ t_{b} = \alpha_0 + \alpha_t r_{e} + \varepsilon_t \] (13)
Model 2: \[ x_t = \beta_0 + \beta_t r_{e} + \varepsilon_t \] (14)
Model 3: \[ m_t = \omega_0 + \omega_t r_{e} + \varepsilon_t \] (15)

Model 1 forms the relationship between trade balance and real exchange rate; Model 2 institutes the link between real exchange rate and volume of exports while Model 3 establishes the link between real exchange rate and the volume of imports. The results of these long-run relationships are presented in Table 3.

Table 3: Long-run model

<table>
<thead>
<tr>
<th>Equation/Model Variable</th>
<th>Dependent Variables</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Ln ( t_b )</td>
</tr>
<tr>
<td>Constant</td>
<td>-0.79***</td>
</tr>
<tr>
<td>Ln ( r_{e} )</td>
<td>(-4.25)</td>
</tr>
<tr>
<td></td>
<td>0.22***</td>
</tr>
<tr>
<td>Observation</td>
<td>180</td>
</tr>
<tr>
<td>Adjusted R-square</td>
<td>0.08</td>
</tr>
<tr>
<td>F-statistic</td>
<td>18.69***</td>
</tr>
</tbody>
</table>

Before testing for the existence of asymmetric cointegration, we conduct the Engle-Granger (1987)'s two-step cointegration test; based on equation (7), the results of these tests are presented in Table 4.

Table 4: Engle-Granger ADF Cointegration tests

<table>
<thead>
<tr>
<th>Model</th>
<th>( t )-statistic</th>
<th>1%</th>
<th>5%</th>
<th>10%</th>
</tr>
</thead>
<tbody>
<tr>
<td>( t_{b} )</td>
<td>1</td>
<td>-17.069</td>
<td>-2.578</td>
<td>-1.942</td>
</tr>
<tr>
<td>( x )</td>
<td>2</td>
<td>-4.218</td>
<td>-2.578</td>
<td>-1.942</td>
</tr>
<tr>
<td>( m )</td>
<td>3</td>
<td>-19.759</td>
<td>-2.578</td>
<td>-1.942</td>
</tr>
</tbody>
</table>

The Engle-Granger cointegration test results (Table 4) above reveal the existence of long-run relationships between real exchange rate and trade balance, exports volume & imports volume in models 1, 2 and 3 respectively. With this, therefore, the null hypothesis of no cointegration is hereby rejected at 1% significance levels. Thus there might be a chance to discover an asymmetric cointegration in the model without further prove from Johansen Jusellius (1990) cointegration test. We now explore further analysis called threshold cointegration using TAR and M-TAR for the three models. The results are presented in Table 5.

Table 5: Enders-Siklos asymmetric cointegration tests

<table>
<thead>
<tr>
<th>Model</th>
<th>H(_0): ( \rho_1 = \rho_2 = 0 )</th>
<th>H(_0): ( \rho_1 = \rho_2 )</th>
</tr>
</thead>
<tbody>
<tr>
<td>TAR:</td>
<td>(\Phi)</td>
<td>M-TAR:</td>
</tr>
<tr>
<td>1(( t_{b} ))</td>
<td>7.48**</td>
<td>1</td>
</tr>
<tr>
<td>Model 2(( x ))</td>
<td>14.43***</td>
<td>1</td>
</tr>
<tr>
<td>Model 3(( m ))</td>
<td>12.03***</td>
<td>1</td>
</tr>
</tbody>
</table>
Note: k Notation is the default lag periods of lagged difference term, ** indicates significance at 5% level, the critical values of $\Phi$ and $\Phi^* \text{ statistics are given in Enders and Siklos (2001).}$ $F$ indicates $F$ statistic for the null hypothesis of symmetric adjustment, $\rho_1 = \rho_2$.

In Table 5 above, we have found, using model 1, asymmetric cointegration occurs between trade balance and real exchange rate in both TAR and M-TAR under the joint null hypothesis. Exhausting model 2 and 3, we also found asymmetric cointegration between exports demand, imports demand and real exchange rate respectively, using TAR model as $\Phi$ and M-TAR model as $\Phi^*$ statistic, both significance at 1% and TAR $F$-statistic of Wald coefficients tests is significant at 5% level. All results are significant at 5% and 1% levels.

**Threshold error-correction models**

Having confirmed the existence of asymmetric cointegrations between real exchange rate and all the trade variables used in this study, we now go further to estimate TAR and M-TAR error correction models so as to evaluate the short run dynamically. We set the maximum lag order of first differenced variables to 12. The lag order is to be trimmed down if and only if the last lag is not found significant at least 10% level. Table 5, Table 6 and Table 7 present the results; M-TAR for trade balance model and TAR for export and import models respectively. All the three models are diagnosed with robust evidence.

**Table 6: Momentum Threshold Autoregressive (M-TAR) Error Correction Model for trade balance**

<table>
<thead>
<tr>
<th>Independent Variable</th>
<th>Dependent Variable: (M-TAR) $\Delta tb$</th>
</tr>
</thead>
<tbody>
<tr>
<td>Constant</td>
<td>-0.003</td>
</tr>
<tr>
<td>$M_t \epsilon_{t-1}$</td>
<td>$-0.442^{***}$</td>
</tr>
<tr>
<td>$(1-M_t)\epsilon_{t-1}$</td>
<td>$-0.269^{***}$</td>
</tr>
<tr>
<td>$\Delta tb_{t-1}$</td>
<td>$-0.308^{***}$</td>
</tr>
<tr>
<td>$\Delta tb_{t-2}$</td>
<td>$-0.281^{***}$</td>
</tr>
<tr>
<td>$\Delta tb_{t-9}$</td>
<td>$0.244^{***}$</td>
</tr>
<tr>
<td>$\Delta rer_{t-1}$</td>
<td>$0.023^{*}$</td>
</tr>
<tr>
<td>$\Delta rer_{t-10}$</td>
<td>$-0.026^{**}$</td>
</tr>
</tbody>
</table>

**Adjusted R-square** 0.420

**F-statistics** 18.128***

**Diagnostic tests:**

- **Serial correlation** 0.867
- **Heteroskedasticity** 0.379

**Wald test:**

- **F-statistics** 8.690***
- **Chi-square** 43.45***

Note: ***, ** and * denote 1%, 5% and 10% significance levels respectively.

After the estimate of M-TAR error-correction model of trade balance, we found that long-run relationship exists between real exchange rate and trade balance and with a causal adjustment manner being asymmetric. Indeed, the rate of the adjustment parameters show that as trade balance and real exchange rate set off from their original steady-state temporarily, the movement or adjustment back to the steady state, is faster when there is relative increase in trade balance (above long-run value), compared to relative decrease in trade balance (below long-run value). We explain this scenario in the following models.

$$\Delta tb_t = K - 0.442[\epsilon_{t-1} + 0.79 - 0.22rer],$$

for $\epsilon_{t-1} > 0$ (16)

$$\Delta tb_t = K - 0.269[\epsilon_{t-1} + 0.79 - 0.22rer],$$

for $\epsilon_{t-1} < 0$ (17)

In the above models, $K$ represents constant and lagged changes in equations 11 and 12. The specification of M-TAR disclosed a return to an underlying equilibrium relationship is faster, 44% adjustment speed, as a short-term movement from equilibrium relationship, are instigated by relative increases in the trade balance, or equally, decreases in the real exchange rate. However, the return to the original equilibrium is only 27% as the temporary moves from equilibrium point are triggered by comparative decreases in the trade balance, or equally, increases in the real exchange rate.

From the investigations made, an inference is drawn that there is an evidence of deficit trade balance in Thailand. This can be explained by the fact that the adjustment of trade balance back to the long run value is slower when it faces deficit than when it faces surplus. The apprehension is that a trade deficit that lasts long can lead to more foreign debt as the country pays more than its earnings, and eventually devaluation of the currency. Thailand historically has been having a negative trade balance with the rest of the world for most of the times. If a trade balance surplus is present, it remains relatively small especially the period between 2001 and 2010. Though if not for crude oil imports, the trade balance would have been positive all the time. There was also a noticeable striking trade deficit in the 1995 – 2000 periods, as a result of Asian economic crisis. Prior to the said crisis, Thailand was importing a lot of capital goods; which was a good thing to reckon, however, it is obvious that too much of a good thing turned out to be bad at the time. These trade deficits occurred despite various trade and investment policies adopted by the government of Thailand, such as export promotion policies that have to do with refinancing facilities, investment promotion for exports and tax refunds. These policies date back from the 1960s through to 1990s.
Table 7: Threshold Autoregressive (TAR) Error Correction Model for export demand

<table>
<thead>
<tr>
<th>Independent Variable</th>
<th>Dependent Variable</th>
<th>Equation</th>
</tr>
</thead>
<tbody>
<tr>
<td>Constant</td>
<td>0.003</td>
<td></td>
</tr>
<tr>
<td>$l_t e_{t-1}$</td>
<td>-0.389***</td>
<td>(21)</td>
</tr>
<tr>
<td>$(1-l_t) e_{t-1}$</td>
<td>-0.095**</td>
<td>(20)</td>
</tr>
<tr>
<td>$\Delta x_{t-1}$</td>
<td>-0.215***</td>
<td></td>
</tr>
<tr>
<td>$\Delta x_{t-12}$</td>
<td>0.425***</td>
<td></td>
</tr>
<tr>
<td>$\Delta rer_{t-1}$</td>
<td>1.117***</td>
<td></td>
</tr>
</tbody>
</table>

Adjusted R-square 0.475

F-statistics 31.024***

Diagnostic tests:

Serial correlation 0.4870
Heteroskedasticity 0.4077

Wald test:

F-statistics 30.056***
Chi-square 90.167***

Note: *** and ** denote 1% and 5% significance levels respectively.

From Table 7, we noticed the existence of long-run cointegration between real exchange rate and export volume. We also observed a causal adjustment manner being asymmetric. When export volume and real exchange rate depart temporarily from the state of equilibrium relationship, adjusting back to steady equilibrium state is faster resulting from a comparative increase in export volume (above long-run value) as compared to a relative decrease in export demand which has very low percentage point of 9%. This means that the error-correction export demand model provides a somewhat meaningless evidence for the adjustment of export demand when it is below long-run value. We have the following as an illustration for this non-meaningfulness:

\[
\Delta x_t = K - 0.389[x_{t-1} - 21.21 + 2.31 rer_{t-1}], \\
\text{for } \hat{e}_{t-1} > 0
\]  
(18)

\[
\Delta x_t = K - 0.095[x_{t-1} - 21.21 + 2.31 rer_{t-1}], \\
\text{for } \hat{e}_{t-1} < 0
\]  
(19)

Where $K$ signifies other terms, which is constant and lagged changes in export and real exchange rates. Error-correction coefficients estimate suggest that return to the underlying equilibrium correlation is faster, with 39% adjustment speed as temporary departures from the original equilibrium is caused by a relative increase in export demand. On the other hand, the return to the underlying equilibrium is just as low as 9% speed of adjustment as temporary departures from it are triggered by relative decreases in export volume or consistently, increases in the real exchange rate. This result confirmed persistent evidence of trade balance deficit observed in Table 5.

Table 8: Threshold Autoregressive (TAR) Error Correction Model for import demand

<table>
<thead>
<tr>
<th>Independent Variable</th>
<th>Dependent Variable</th>
<th>Equation</th>
</tr>
</thead>
<tbody>
<tr>
<td>Constant</td>
<td>0.013**</td>
<td></td>
</tr>
<tr>
<td>$l_t e_{t-1}$</td>
<td>-0.095**</td>
<td>(21)</td>
</tr>
<tr>
<td>$(1-l_t) e_{t-1}$</td>
<td>-0.385***</td>
<td>(20)</td>
</tr>
<tr>
<td>$\Delta m_{t-1}$</td>
<td>-0.420***</td>
<td></td>
</tr>
<tr>
<td>$\Delta m_{t-7}$</td>
<td>-0.118*</td>
<td></td>
</tr>
<tr>
<td>$\Delta m_{t-9}$</td>
<td>0.119*</td>
<td></td>
</tr>
<tr>
<td>$\Delta m_{t-11}$</td>
<td>-0.154**</td>
<td></td>
</tr>
<tr>
<td>$\Delta rer_{t-3}$</td>
<td>-0.724*</td>
<td></td>
</tr>
</tbody>
</table>

Adjusted R-square 0.332

F-statistics 12.873***

Diagnostic tests:

Serial correlation 0.992
Heteroskedasticity 0.167

Wald test:

F-statistics 9.548***
Chi-square 47.74***

Note: *** denotes 1%, 5% and 10% significance levels respectively.

In Table 8, we have import demand model for TAR error-correction, the model also advocates speedy adjustment of import demand, and however, the adjustment is faster when below long-run value. On the other hand, the model provides very little evidence concerning adjustment in import demand when it is above long-run value. This can be demonstrated as follows:

\[
\Delta m_t = K - 0.095[m_{t-1} - 22.00 + 2.54 rer_{t-1}], \\
\text{for } \hat{e}_{t-1} > 0
\]  
(20)

\[
\Delta m_t = K - 0.385[m_{t-1} - 22.00 + 2.54 rer_{t-1}], \\
\text{for } \hat{e}_{t-1} < 0
\]  
(21)

Where $K$ expresses other terms, which is lagged changes and constant in imports and real exchange rates. The assessed coefficient of error correction recommends that 38% deviation of imports in last-period from its long-run value will be corrected by imports change. Specifically, the values of the conformity parameter showed that when real exchange rate and imports coincidentally move from their underlying equilibrium relationship, alteration back to balance is faster after a relative reduction in imports or similarly, an increment in the real exchange rate. On the other hand, imports adjustment is found to be less meaningful when it is above its long-run value. This result is in opposite with the one in Table 6, and it shows that a shock of exchange rate on import volume is possible to be temporary in nature. To be precise, the increment in exchange rate has the same impact on both exports and imports demand in Thailand.
We have also included a control variable of political stability in our short-run model, using autoregressive distributed lag (ARDL), in order to capture its effects on the trade balance in Thailand. We found that, although, the variable has a negative effect the effect is not significant.

Table 9: Short-Run Model

<table>
<thead>
<tr>
<th>Regressors</th>
<th>Coefficient</th>
<th>Standard Error</th>
<th>T-Ratio</th>
</tr>
</thead>
<tbody>
<tr>
<td>DLEXR</td>
<td>-0.0259**</td>
<td>0.01109</td>
<td>-2.3417</td>
</tr>
<tr>
<td>DLXVOL</td>
<td>0.3770***</td>
<td>0.07061</td>
<td>5.3553</td>
</tr>
<tr>
<td>DLMVOL</td>
<td>-0.3973***</td>
<td>0.07124</td>
<td>-5.5779</td>
</tr>
<tr>
<td>DPSTAB</td>
<td>-0.4410</td>
<td>0.64620</td>
<td>-0.6823</td>
</tr>
<tr>
<td>Dependent Variable DTB, R-Squared 0.99578, DW-statistic 1.9979</td>
<td></td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

Robustness Checks

It is imperative to have an additional support to our findings. In doing this, we conduct two types of estimation; the dynamic ordinary least square (OLS) and the fully modified ordinary least squares (FMOLS), to serve as robust checks to our findings. We presented the results of the robustness checks in Table 10 with trade balance as the dependent variable. The coefficients of exchange rate using DOLS is negative and significant at 10 percent which is in line with our main findings. The coefficient is also negative and significant at 5 percent using FMOLS. This confirms our findings that exchange rate causes negative trade balance in Thailand, as a result of excess import over export. The other two independent variables; import and export volumes showed a negative and a positive relationship with trade balance respectively and are significant at 1 percent level. Political stability showed a non-significant negative effect on trade balance using both FMOLS and DOLS.

Table 10: Results for DOLS and FMOLS Estimates

<table>
<thead>
<tr>
<th>Dep. Var. : Trade Balance</th>
<th>DOLS</th>
<th>FMOLS</th>
</tr>
</thead>
<tbody>
<tr>
<td>Exchange Rate</td>
<td>-0.12*</td>
<td>-0.022*</td>
</tr>
<tr>
<td></td>
<td>(0.06)</td>
<td>(0.01)</td>
</tr>
<tr>
<td>Import Volume</td>
<td>-1.00***</td>
<td>-1.00***</td>
</tr>
<tr>
<td></td>
<td>(0.01)</td>
<td>(0.01)</td>
</tr>
<tr>
<td>Export Volume</td>
<td>0.99***</td>
<td>0.99***</td>
</tr>
<tr>
<td></td>
<td>(0.01)</td>
<td>(0.01)</td>
</tr>
<tr>
<td>Political Stability</td>
<td>-0.001</td>
<td>-0.004</td>
</tr>
<tr>
<td></td>
<td>(0.00)</td>
<td>(0.00)</td>
</tr>
</tbody>
</table>

Note: values in parenthesis are standard errors

CONCLUSION

In examining the impacts of exchange rate shock on trade variables, most studies currently employed asymmetric threshold cointegration framework. However, prior to this, several studies commonly used symmetric cointegration analysis to examine the interdependence among variables. One of such methods is implemented from Engle-Granger (1987), which assumes symmetric adjustment of the error correction term. For symmetric cointegration, adjustment coefficients are alike, irrespective of their positivity or negativity in the equilibrium error. This study tries to investigate the long run equilibrium relationship between real exchange rate and trade balance, export demand as well as import demand using asymmetric cointegration framework advanced by Enders and Siklos (2001). The study also incorporates political stability in its short run model as a control variable, to see whether the variable has any effect on Thailand’s trade balance. This method is used because; it has been, nowadays, an important process in macroeconomic analysis, with quite a number of studies providing clear evidence of asymmetric adjustment of macroeconomic variables.

In this study, we found that there exists a long run asymmetric cointegration between trade balance and real exchange rate when we conducted the M-TAR model. Using TAR model, we also found the existence of asymmetric cointegration between real exchange rate and exports and imports volumes with the value of threshold not equal to zero (\( r \neq 0 \)). Therefore, based on the expectations and objective of this study, we report that changes in the real exchange rate have contributed to the recent trade balance deficit came across by Thailand.

From the M-TAR model of the error correction for trade balance, an inference can be drawn that long-run relationship exists between trade balance and real exchange rate with the underlying adjustment manner being asymmetric. Basically, the value of the adjustment parameters specifies that when trade balance and real exchange rate depart from their core equilibrium relationship, then adjustment back to equilibrium is more speedy resulting from an increase in trade balance (above long-run value) as related to decreasing in trade balance (below long-run value). The export model for TAR error correction suggests that when export volume and real exchange rate momentarily depart from their equilibrium correlation, adjustment back to original equilibrium is more rapid resulting from a relative increase in export volume (above long-run value) compared to a relative decrease in export volume (below long-run value).

As per import demand model for TAR error correction, the model recommends rapid adjustment of import volume when it is below long-run value. The model also asserts that returns or adjustment back to equilibrium point is quite slow when it is above long-run value and faster when it is below long-run value. The results indicate the evidence that there is the presence of fluctuations of trade balance deficit and surplus in Thailand, with much effect on trade balance deficit. This is because adjustment of trade balance back to its long-run value is quite slow when it
faces deficit than when it faces surplus. The study also uses political stability in its short run model as part of the explanatory variables to capture its effects on trade balance; however, its effect on trade balance was not a meaningful one. For the trade balance deficit, it is most likely due to a policy of massive crude oil imports between late 1990's through to 2010 which in turn led to the devaluation of Thailand’s currency. The shock of the exchange rate, import demand might be more or less a temporary one. Therefore, to avoid a negative effect on Thailand’s currency, it is suggested that policy adjustment is implemented to boost export and in addition reduce over-reliance on crude oil imports.

REFERENCES


The Bank of Thailand, Report Published 2015.


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