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South East Asian Monetary Integration: New Evidences from Fractional Cointegration of Real Exchange Rates

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Abstract

We study the long-run relationship of real exchanges rates (RERs) among the ASEAN-5 countries by testing the theory of Generalized Purchasing Power Parity (G-PPP) from the new perspective of fractional cointegration. The long-run co-movements of the RERs are examined by applying a recent estimator of fractional cointegration that consists of a frequency Whittle approximation of the cointegrating system's likelihood function. The contribution of the fractional cointegration study is justified by identifying several weak fractional cointegration relationships that signal that deviations of RERs from their long-run equilibrium are highly persistent. These findings contrast with all previous studies that restrict their investigation to the traditional $I(1)/I(0)$ cointegration. Our results support further monetary integration among different sub-groups of the ASEAN-5 countries as they share long-run comovements with each others. However, a full-fledged monetary union embracing all ASEAN-5 members is still limited from the perspective of the G-PPP theory.

Keywords: Monetary Union, Fractional Cointegration, Generalized purchasing power parity, ASEAN
JEL: F31, F33

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1. Introduction

This empirical paper examines the long-run relationship between the real exchange rates (RERs) of the five most advanced members of the ASEAN countries [3]. All the literature based on the cointegration of RERs is concentrated on the traditional $I(1)/I(0)$ cointegrating relationships. In the former case, there is no equilibrium relationship and related studies conclude against the existence of an Optimal Currency Area (OCA, see Sun and Simons (2011) ; Wilson and Choy (2007)), while in the latter case, all RERs share the same kind of real disturbances, thus strengthening the case of a currency union (Mishra and Sharma (2010) ; Ogawa and Kawasaki (2008)). However, classical cointegration models used in these studies imply short memory residuals as the equilibrium errors are restricted to an I(0) process.

Our study examines the emergence of an OCA in the ASEAN-5 countries, thus extending a literature for which there is no consensus. Considering limitations induced by traditional cointegration, we suppose that the shocks on the ASEAN-5 countries' RERs can be extremely persistent, thereby implying the presence of long memory in the residuals of the cointegrating relationship and, accordingly, a slow return towards the long-run equilibrium. Conversely, in the case of strong RERS's long-run relationships, the cointegrating residuals could respond more quickly to shocks involving less persistent deviations from equilibrium. Fractional cointegration analyses allow us to capture these relationships by considering that residuals can follow a fractional process $I(d)$.

The issue of residuals with respect to the fractional process has important implications for assessing the regional convergence of the fundamentals that drive the evolution of the RERs. Accordingly, the extent to which RERs are cointegrated is relevant information for analyzing more precisely the degree of monetary integration in the ASEAN-5 countries. The existence of possible misalignments among the relative prices in regional goods markets can pose serious obstacles to the development of the intra-regional trade, thus slowing the integration process. As trade is the main driver of Asian economic integration, it is crucial for these countries to reach a large zone of price stability. In this regard, the choice of an appropriate monetary anchor as an informal form of policy coordination must be determined before moving on to more advanced forms of monetary integration, such as a full-fledged monetary union. Considering the importance of both currencies inside the region, we test the G-PPP hypothesis introduced by Enders

[3] The ASEAN-5 consists of founder members namely Singapore, Thailand, the Philippines, Indonesia, Malaysia.

and Hurns (1994) by envisaging both dollar-based and yen-based RERs for the ASEAN-5 countries [4]. Moreover, we implement a third numeraire in our empirical strategy defined as a basket currency that is equally weighted following the literature on basket-based arrangements for East Asian countries.

According to Enders and Hurns (1994), in a currency area, the forcing variables will be sufficiently integrated for the RERs to share common trends. In this circumstance, each real income process will share common disturbances, and RERs between any two countries will therefore be cointegrated. Thus, if macroeconomic fundamentals, which drive RERs in the ASEAN-5 countries, such as the business cycles, expenditures patterns or money supply are highly interrelated, RERs may share a common equilibrium in the long-run that supports further monetary integration among the countries.

The main results of the paper are the following: 1) We find two different sub-groups of the ASEAN-5 countries sharing with each other a long-run relationship. The first composed of Indonesia, Malaysia, the Philippines, and the second composed of Malaysia, the Philippines and Thailand. 2) All of the cointegration relationships are weak in the sense that RERs can deviate persistently from the long run equilibrium with an evident tendency for reversion towards this equilibrium. As expected, these two results imply important limitations for an immediate monetary union involving all the ASEAN-5 countries from the perspective of the G-PPP theory. These findings suggest that economic structures lack homogeneity in terms of monetary and real linkages, which results in significant misalignments of the RERs despite the fact that trade in goods and services and foreign direct investment (FDI) activities have expanded rapidly over the last three decades. 3) All significant cointegration relationships exhibit a positive long run coefficient, thus implying that countries that share a long-run equilibrium adjust in a symmetric way to macroeconomic disturbances. 4) Both the yen and the US dollar are important when considering a step-by-step approach towards the coordination of monetary policy. 5) There are some evidences to assert that the ASEAN-5 countries became more integrated after 1997-98 as we find more cointegrated pairs in the post-crisis subsample. More importantly, the equilibrium errors respond more rapidly to shocks, thus resulting in less persistent deviations towards the common equilibrium. This result suggests that structural reforms in the aftermath of the crisis

[4] Indeed, the US is one of the most important export markets for final outputs produced inside the region through vertical integration of production chains, while the Japanese economy accounts for a large proportion of trade among Southeast Asian countries

appear to have resulted in increased long-run relationships of RERs.

The remainder of the paper is organized as follows. Section 2 provides an overview of previous studies and clarifies the position of our analysis. Section 3 presents the data and introduces our empirical methodology of estimation. In section 4, we discuss the results and their economic implications. Section 5 concludes the paper.

2. Literature review

2.1. Monetary integration in East Asia

The economic regionalism in East Asia received political momentum after the Asian financial crisis prompting regional authorities to establish preventive mechanisms with the aim of strengthening monetary and financial cooperation. Furthermore, the economic regionalism led to substantial reforms[5] such a regional economic surveillance process framed in the Economic Review and Policy Dialogue process (ERPD, May 2000) and a liquidity support arrangement under the form of bilateral swaps (Chiang Mai Initiative, May 2000). Policy dialogue on macroeconomic and financial issues are essentially conducted through the ASEAN+3[6] considered as the favored political framework in order to pursue the regional process. Recognizing the close economic interdependence within the region and the need to promote regional trade and growth, Asian authorities stated their intention to create a common single currency in the future[7]. Arguments in favor of monetary integration are commonly related to the preponderance of market-driven force, such as trade and FDI (Rana (2007)). Their rapid increase in volume are mainly explained by the international segmentation of production processes (the so-called global production sharing), which led to the constitution of an intense intra-industry trade networking. According to Shin and Wang (2003) and Lee and Azali (2010), the case for a monetary union in East Asia constitutes an option that policymakers should envisage in the near future, considering the recent trend

[5]See, for instance, (Kawai (2010)) for a comprehensive survey of recent development on financial and monetary issues.

[6]The ASEAN+3 is a forum began in 1997 and is composed of Singapore, Thailand, the Philippines, Indonesia, Malaysia, Vietnam, Brunei, Laos, Cambodia, Burma as members of the ASEAN, plus the three East Asian nations of China, Japan, and South Korea

[7]At the 39th Annual Meeting of the ADB, a first step was initiated in this direction with the introduction of the Asian currency unit (ACU) as a benchmark to monitor movements in the values of currencies in the region. Moreover, the prospect of launching a single currency was recently put forward by the Japanese prime minister on 23 October 2009 during the 15th summit of the Association of South-East Asian Nations.

in intraregional trade flows and the critical role of intra-industry trade in the regional integration process.

Ordinarily, fixed exchange rate arrangements refer to the adoption of one single currency as a definitive or transitory nominal anchor to synchronize currency adjustment prior to the creation of a new currency. Some studies have suggested that the US dollar will be better suited for East Asian countries to guarantee implicit coordination of exchange rates movements and then anchor domestic price levels within the region (McKinnon and Schnabl (2004)). This argument is generally rationalized, given that East Asian countries have returned after the 1997-98 crisis to the pre-crisis level of the dollar pegging at high frequencies and thus creating a *de facto* dollar bloc. The creation of a yen block has also been envisaged (Kwan (2001)) following the announcement of Japanese authorities to promote the yen's internationalization while the possibility of a common currency area based on the Chinese yuan (Shirono (2009), Park (2010)) is still limited. Moreover, considering the intrinsic vulnerability of a unilateral peg for countries with a diversified trade pattern (Bird and Rajan (2002), Rajan (2002), Wilson et al. (2007)) and coordination failure in choosing an appropriate regional arrangement (Ito et al. (1998), Ogawa and Ito (2002), Bénassy-Quéré (1999)), the idea of a weighted basket peg composed of international currencies (Williamson (2005)) or even regional currencies (Ogawa and Shimizu (2006), Watanabe and Ogura (2010)) is gaining momentum.

2.2. Are the shocks facing the countries correlated?

From a theoretical perspective, monetary unions are studied through the lens of the OCA theory. Since Mundell (1961), different economic criteria have been established as preconditions to achieve monetary integration among a group of countries planning to adopt a common monetary standard. As a monetary policy cannot be tailored to absorb country specific shocks within a monetary union, two countries may find an advantage in adopting the same currency if there are alternative mechanisms of adjustments that can help them to cope with asymmetric shocks. The seminal paper of Mundell (1961) argued that the mobility of the labor force within a union and the flexibility of prices/wages could be considered as alternatives to flexible exchange rates to absorb asymmetric shocks[8]. However, if countries face symmetric shocks, then the need for alternative mechanisms is considerably reduced. Therefore, the synchronization of business cycles

[8] See, also, McKinnon (1963), Kenen (1969) and De Grauwe (2005) for a survey of the OCA theory.

and the symmetry of output shocks are typically considered as being one of the most important criteria from the perspective of OCA theory. However, literature is inconclusive on this point as empirical results remain particularly affected by the sample period and the econometric tools considered.

A frequently used approach to assess the correlation of shocks within a group of countries is the methodology proposed by Blanchard and Quah (1989), which allows for the distinction between aggregate demand and supply disturbance[9]. In related studies, Bayoumi and Eichengreen (1994) find that supply shocks are symmetric, on the one hand, between Singapore, Malaysia, Indonesia and Hong-Kong; on the other hand, between Japan, Taiwan and South Korea. Chow and Kim (2003) found that domestic outputs of East Asian countries are strongly influenced by country-specific shocks while regional shocks are far more important in European countries that have joined the Economic and Monetary Union. Considering a time-varying analysis, Zhang and Sato (2008) suggest that the greater China economies (Mainland China, Hong Kong and Taiwan) are good candidates for a monetary union. Obiyathulla (2008) examines the feasibility of a Common Currency Area (CCA) for 14 East Asian countries and finds that while a region-wide CCA is not feasible, several paired clusters are potential candidates. Additionally, though Genberg and Siklos (2010) do not clearly identify a group of countries for which shocks are unambiguously highly correlated, Ahn et al. (2006) find a group of seven East Asian economies (Indonesia, Malaysia, Singapore, Thailand, Hong Kong SAR, Korea, and Taiwan) qualified for an OCA in terms of macroeconomic shocks. As can be observed, these results do not lead to firm conclusions[10].

Different empirical methods have been applied by Girardin (2004) and Girardin (2005) who uses Markov-switching techniques, assuming that synchronization of East Asian growth cycles are regime-dependent. His findings do not support a monetary arrangement involving the yen, as East Asian countries are not fully synchronized with Japan. Moreover, according to the author, China would be a better candidate for monetary integration with East Asian countries with Japan. Allegret and Essaadi (2011) introduce a spectral analysis based on the time-varying coherence function and find the presence of a common cycle in East Asia after the crisis of 1997-98, thus suggesting that East Asian countries constitute an OCA. Moneta and Rüffer (2009) find

[9] From a theoretical perspective, demand-shocks have only a transitory impact on the long-term product while aggregate supply-shocks influence the output in a permanent way.

[10] See, also, Kim (2007), Ng (2002), Zhang et al. (2004), Huang and Guo (2006) and Kwack (2004).

a significant common growth dynamic inside the region by using a state-space framework where common movement is captured by unobservable variables influencing the evolution of the GDP growths. Using cointegration techniques, Sato and Zhang (2006) find that some pair-countries in the region share both the long-run and the short-run synchronous movements of the real outputs, particularly among the ASEAN economies consisting of Singapore, Thailand and Indonesia, and among the Northeast Asian region, which consists of Hong Kong, Korea, Mainland China,Japan and Taiwan while Sato et al. (2009) find that ASEAN countries alone are not a feasible group among which to form a monetary union unless Japan is included.

2.3. Cointegration analyses

According to Ogawa and Kawasaki (2008), a group of countries may constitute an OCA if one of the conditions is satisfied although the shocks are simultaneously asymmetrical within this group. Therefore, the ”shocks” literature based on the structural vector autoregression model (SVAR) does not necessarily tell the whole story about the relevance of an OCA within a region. Indeed, the SVAR model includes first-differenced variables, thus removing any information about the long-run equilibrium relationship for a set of real variables. This can pose a greater obstacle for monetary integration if macroeconomic fundamentals among the economies, such as real output variables, are not cointegrated. Accordingly, this can lead to a different growth path for each country in the long-run (Sato et al. (2009)).

A convenient complementary approach to examine the suitability of a monetary union is the concept of G-PPP, as proposed by Enders and Hurns (1994). The G-PPP model was originally suggested as a solution to overcome the empirical failure of the Purchasing Power Parity (PPP) in adequately explaining RER behavior[11]. According to the authors, the failing of the PPP is mainly due to the continuous effect of real and nominal shocks on variables that drive RERs and, as consequence, imply permanent movement in RERs. The intuition for this result is rather straightforward. If macroeconomic fundamentals that determine RERs, such as real output, expenditure patterns and money supply, are, themselves, nonstationary, then RERs are also nonsta-

[11] PPP theory is a structural model of RER determination supposing that two currencies in equilibrium lead to the same purchasing power in both countries. From a empirical perspective, PPP holds if bilateral RER is stationary involving RER convergence to a level in the long run. However, numerous empirical studies have shown that RERs are non-stationary in level, thereby calling into question the validity of PPP as a structural model of RER determination (Alder and Lehman (1983), Enders (1988)). In the context of Asian countries, see Choi et al. (2011), Kim et al. (2009), Liew et al. (2004) and Wu et al. (2004).

tionary. This finding is consistent with many structural models of exchange rate determination including Dornbush (1976) and Frankel (1983).

Enders and Hurns (1997) suggest that “despite the fact that domestic economies differ, the trends in real forcing variables affecting real exchange rates might be similar across diverse countries” in such a way that RERs may share marked relationships with each other. Indeed, PPP deviations in the long run might imply that non-stationary RERs move together consistent with common trends in both domestic and international environments, thereby supporting the existence of a long run equilibrium relationship between RERs. From the perspective of the OCA theory, this mean that a shock in any one RER within a group of two or more countries will affect other RERs, thus supporting the case of an OCA in the sense of Mundell (1961). This group would appear sufficiently interrelated and would, therefore, meet preconditions to adopt a common single currency.

As the long-run relationship among regional currencies identified by the G-PPP model can be regarded as a precondition of an OCA, several recent papers have proposed examining whether the East Asian currencies shared a common trend in the long-run. Using real exchange rates of the ASEAN-5 countries, the empirical results of Wilson and Choy (2007) do not support the evidence of a currency area with either the US or Japan, while Mishra and Sharma (2010), citing results robust enough to change the choice of a base currency, find the opposite for seven ASEAN countries and India. Rangkakulnuwat et al. (2010) extend the theory of G-PPP by developing a model including foreign variables such as the real money supply and the output and interest rate. They further found that selected Asian countries form a single currency area with Japan as the base country. Choudhry (2005) conducted Johansen's cointegration test and found a substantial change in relationships between several RERs of the Far East Asian countries with ample evidence of G-PPP after the crisis period. Sun and Simons (2011) find a ”pentagon” group of five countries (South Korea, the Philippines, Thailand, Malaysia and Indonesia) with potential success for further monetary integration even if neither a yen bloc nor a US dollar bloc is forming in East Asia. Ogawa and Kawasaki (2008) investigate the possibilities of an OCA in East Asia and find that the ASEAN+3 countries are forming an OCA, notably in the post-crisis period, thus providing the possibility of adopting a common exchange rate policy with reference to a common currency basket composed of the US dollar, the euro and the yen.

However, these papers neglect the fact that RERs probably converge slowly towards a long-run equilibrium, implying that the autocovariance function of the cointegrating residuals declines at a slower hyperbolic rate rather than the classical geometric rate found in the traditional cointegrating residuals (see the next section). To overcome this conceptual drawback, we propose a new empirical methodology based on fractional cointegration.

3. Data and empirical methodology

3.1. Description of the data

All monthly data include the consumer price index (CPI) and the nominal bilateral exchange rates against two numeraire, the US dollar and the Japanese yen, for five East Asian countries, namely, Malaysia (Mala.), Indonesia (Indo.), Singapore (Sing.), Philippines (Phil.), Korea and Thailand (Thai.). The bilateral RERs against the dollar (USD) and the yen (YEN) are calculated according to the following formulas:

$$q_t = s_t + p_t^* - p_t$$

where q_t is the bilateral RER, s_t is the nominal exchange rate (expressed as the domestic currency price of the base currency), and p_t^* and p_t are the base country price indices and the domestic price index respectively. We also compute a third numeraire constructed from dollar-based and yen-based RERs, as a basket composed of both currencies equally weighted (BSK). All data were collected from international monetary fund *International Financial Statistics* Database between 1975:1 and 2011:12. For estimation, we take the natural logarithms of all variables normalized to 0 at the initial observation 1975:1.

The sample is divided into two sub-samples, the pre-crisis period (1975:1-1997:5) and the post-crisis period (1998:6-2011:12), because the crisis may have permanently affected some macroeconomic fundamentals of East Asian economies. Furthermore, long run relationships may have been significantly modified in such a way that a currency union may be more or less plausible after the crisis. This Investigation is also performed on the full sample.

3.2. Strategy of estimation and testing

In this paper, we conduct our analysis using fractional cointegration, that is, a generalization of the traditional $I(1)/I(0)$ cointegration in the sense that integration orders of the regressor (δ)

and the cointegrating residuals (γ) are real numbers and not just integers. In addition to being less restrictive, the fractional cointegration can take into account weak fractional cointegration relationships ($\delta - \gamma < 1/2$) that escape with the traditional cointegration estimators that are based on the existence of a very strong reduction of the dimension. The estimations of bivariate fractional cointegration models usually involve in two steps. The first step is to estimate the long-run coefficient (β), and the second step is to estimate the long memory parameter of the cointegrating residuals. This methodology has the advantage of being easy to use though it requires an assumption on the order of integration of the regressor ($\delta = 1$) and induces a loss of efficiency due to the two-step estimation. Therefore we use a parametric estimator introduced by de Truchis (2012) and that operates in one step and allows us to simultaneously estimate the orders of integration of the regressors and residuals in addition to the correlation long run coefficient. We consider the following cointegration model:

$$(1 - L)^{\gamma}(y_t - \beta x_t) = \varepsilon_{1t}, \qquad (1 - L)^{\delta} x_t = \varepsilon_{2t}, \qquad t = 1, 2, ..., n, \tag{1}$$

where y_t and x_t are two different exchange rates and $(1 - L)^{\alpha}$ is the fractional filter, which is further denoted Δ^{α} and defined by its binomial expansion

$$(1 - L)^{\alpha} = \sum_{k=0}^{+\infty} a_k(\alpha) L^k, \qquad a_k(\alpha) = \frac{\Gamma(k - \alpha)}{\Gamma(k + 1)\Gamma(-\alpha)} \tag{2}$$

$$\Gamma(z) = \int_0^{+\infty} t^{z-1} e^{-t} dt, \tag{3}$$

In equation (1), y_t is said to be cointegrated if the error term $v_t = y_t - \beta x_t$ satisfies $v_t \sim I(\gamma)$ with $\gamma < \delta$ and $x_t \sim I(\delta)$, thus indicating that y_t cannot diverge indefinitely from x_t. Different types of cointegration can be identified depending on δ, γ and the gap $\delta - \gamma$ or, more intuitively, depending on how the equilibrium errors respond to the shocks. We denote strong fractional cointegration when $\delta - \gamma > 1/2$ because it results in a weak persistency of the deviation from long run equilibrium. This includes the traditional $I(1)/I(0)$ cointegration case that involves a very strong return mechanism. Accordingly, we call weak fractional cointegration the case where $\delta - \gamma < 1/2$ with $\delta > 1/2$ because it results in highly persistent deviations from equilibrium due to the non-stationarity of the cointegrating errors. Finally, we name stationary fractional cointegration a relationship where $\delta - \gamma < 1/2$ with $\delta < 1/2$, because it results in a weak dimensionality reduction and stationary cointegration errors. The estimator that we use is has good finite sample properties considering the stationary and the non-stationary regions of δ and γ and consequently, allows us to take into account all the aforementioned cases of cointegration [12]. Exploiting the VAR structure of the cointegration system in (1) and allowing for autocorrelation we obtain,

$$\begin{pmatrix} \varphi_{11}(L) & \varphi_{12}(L) \\ \varphi_{21}(L) & \varphi_{22}(L) \end{pmatrix} \begin{pmatrix} (1-L)^{\gamma} & \\ & (1-L)^{\delta} \end{pmatrix} \begin{pmatrix} \nu_t \\ x_t \end{pmatrix} = \begin{pmatrix} \varepsilon_{1t} \\ \varepsilon_{2t} \end{pmatrix} \tag{4}$$

where $\varphi_{ij}(L)$ denote the lag polynomial of order p, and $\varphi_{ij}(L) = \sum_{k=1}^{p} 1 - \phi_{ij}(k)L^k$ and has all its eigenvalues are less than 1 in modulus. The estimation of (4) operates in the frequency domain and consists of a Whittle approximation of the likelihood function,

$$\mathcal{Q}_n(\theta) = -\sum_{j=1}^{n} \left[\log\det f(\lambda_j) + \operatorname{tr} f(\lambda_j)^{-1} I(\lambda_j) \right], \qquad \theta = (\delta, \gamma, \beta) \tag{5}$$

where $\lambda_j = (2\pi j/n)$ is the angular frequency and $I(\lambda_j)$ refers to the periodogram matrix

$$I(\lambda_j) = \begin{pmatrix} I_{yy}(\lambda_j) - 2\beta I_{xy}(\lambda_j) + \beta^2 I_{xx}(\lambda_j) & I_{xy}(\lambda_j) - \beta I_{xx}(\lambda_j) \\ I_{xy}(\lambda_j) - \beta I_{xx}(\lambda_j) & I_{xx}(\lambda_j) \end{pmatrix}. \tag{6}$$

The cross-periodogram of a_t and b_t, denoted as I_{ab}, is defined as

$$I_{ab}(\lambda_j) = w_a(\lambda_j)\overline{w}_b(\lambda_j), \qquad w_a(\lambda_j) = \frac{1}{2\pi n}\sum_{t=1}^{n} a_t e^{it\lambda_j} \qquad \overline{w}_b(\lambda_j) = \frac{1}{2\pi n}\sum_{t=1}^{n} b_t e^{-it\lambda_j}. \tag{7}$$

The spectral density matrix, $f(\lambda)$, is defined by

$$f(\lambda) = \frac{1}{2\pi}\Lambda(e^{ij\lambda})^{-1}\Phi(e^{ij\lambda})^{-1}\Sigma(\Phi(e^{ij\lambda})^{-1})^{*}(\Lambda(e^{ij\lambda})^{-1})^{*} \tag{8}$$

where Σ is the covariance matrix and Φ is the spectral correlation matrix. We constrain $\varphi_{12}(L) = 0$ to preserve the weak exogeneity of x_t. Finally, parameters δ and γ enter the likelihood function through the matrix Λ:

[12] A Monte Carlo simulation document the finite sample properties of this estimator in the additional content.

$$\Lambda(e^{ij\lambda}) = \begin{pmatrix} |1-e^{ij\lambda}|^{\gamma} & 0 \\ 0 & |1-e^{ij\lambda}|^{\delta} \end{pmatrix} \tag{9}$$

To confirm the use of G-PPP theory, we are also interested in testing for a mean reverting process in the RERs. In other words, we want to test for a fractionally integrated process against a unitary process. We perform the "LV" fractional Wald test proposed by Lobato and Velasco (2007), which consists of restating $(1-L)^{\delta}x_t = \varepsilon_{2t}$ to test that δ is not significantly different from 1. In other words, we test x_t as a unit root process against the fractional alternative:

$$(1-L)x_t = \varphi\zeta_{t-1}(\delta) + \varepsilon_t, \qquad \zeta_{t-1}(\delta) = \frac{(1-L)^{\delta-1}-1}{1-\delta}(1-L)x_t \tag{10}$$

Estimating (10) requires that we pre-estimate δ. Lobato and Velasco (2007) suggest to apply the semi-parametric local Whittle estimator (LWE) investigated by Robinson (1995). Then, using $\hat{\delta}_{LWE}$ in (10), it is obvious that under the null hypothesis, φ is not significantly different from 0 when x_t possesses a unit root. Giving that the LWE is semi-parametric, we must be careful in the selection of bandwidths because of the possible presence of short run dynamics. We have performed the LV test considering different bandwidths ($m = \{0.5, 0.6, 0.7\}$), and no significant changes have been observed [13]. Table 1 reports the results of the fractional unit root test and highlights that all of the examined exchange rates possess a unit root. These results are consistent with previous cited papers that claim that RERs are non-stationary and have no mechanism of mean reversion, thus showing that the long-run PPP does not hold. This finding encourages us to test for the theory of G-PPP.

Obviously, we are also interested in testing for cointegration, that is, testing $\delta - \gamma > 0$. The results (see table 1) suggest that the LV test can be employed using the cointegrating residuals because all exchange rate possess a unit root that amounts to testing for $1-\gamma > 0$. This procedure is suggested in the conclusion of Lobato and Velasco (2007).

[13] Unreported results are available upon request. All computations are performed using RATS v 8.10.

Table 1: LV test based on the LWE ($m = 0.5$)

Num.	x_t	LWE	Stat.	Signif.
USD	Indo.	0.96204	0.19426	0.84597
	Mala.	1.00102	-0.69610	0.48636
	Phil.	1.01946	0.58509	0.55849
	Sing.	1.20625	-0.13485	0.89273
	Thai.	0.97949	0.15841	0.87413
YEN	Indo.	1.00610	0.34554	0.72969
	Mala.	0.95049	-0.09797	0.92196
	Phil.	0.89161	-0.11633	0.90739
	Sing.	0.96317	-0.17088	0.86432
	Thai.	0.97190	-0.49618	0.61977
BSK	Indo.	0.99233	-0.13958	0.88899
	Mala.	0.96478	0.19718	0.84369
	Phil.	0.92032	-0.16490	0.86902
	Sing.	1.03042	-0.08608	0.93140
	Thai.	0.97712	-0.31426	0.75332

4. Empirical results

4.1. Global implication of the G-PPP

The empirical study conducted in this section is based on the examination of bivariate fractional cointegration relationships for 25 pairs of East Asian RERs considering three numeraire: USD, YEN and BSK. The fractional cointegration relationships are tested using the LV test with significance thresholds set at 1%, 5%, and 10%. The results are presented, by country, in tables 2 to 6 for the full sample and the two sub-samples. First, we discuss estimation results obtained from the full sample. The hypothesis of no fractional cointegration is rejected by the LV test in 10 out of 25 pairs when the base currency is the US dollar. Concerning the basket-based RERs, a more important number of significant long run relationships are detected (see figure 1). On the whole, this seems to suggest that a unilateral dollar-based currency union appears to be less suited than for the basket alternative. This finding is supported by the cointegration relationships detected using the yen as the numeraire, where 13 pairs appear. However, the significance of the long-run relationships are mostly invariant to the base currency. Regarding the estimation results based on the currency basket, a well-defined group composed of Thailand, Malaysia and the Philippines exhibit long-term co-movements. This is also the case for Indonesia and the

Philippines. However, these groups breaks up when we consider the U.S. dollar as currency. In contrast, Singapore has only the Philippines as a link to other countries. Interstingly, Indonesia Philippines and Malaysia are fractionally cointegrated with each other thus constituting a second group sharing a common long-run equilibrium but only when the yen is the base currency.

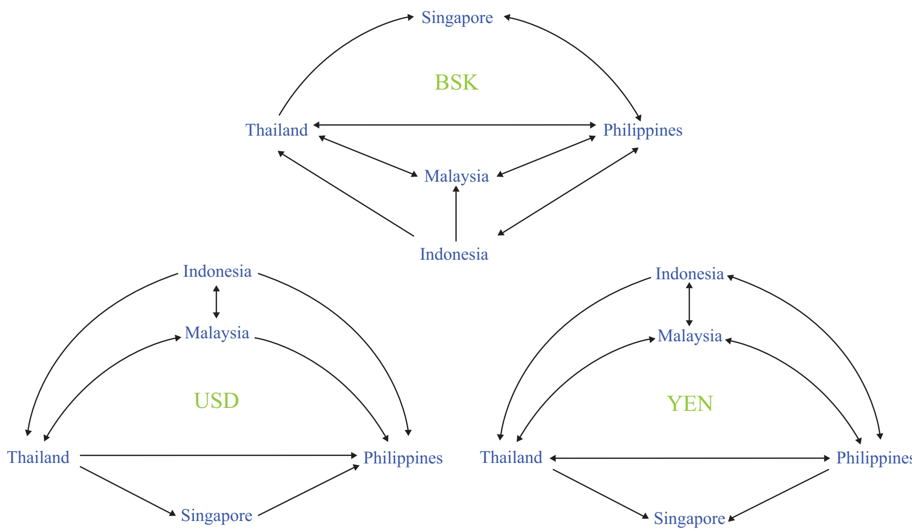

Figure 1: Basket-based, dollar-based and yen-based RER fractional cointegration relationships (full sample analysis).

These results have several implications regarding the existence of an OCA in the ASEAN-5 countries. The most important is that a monetary union involving all ASEAN-5 countries is actually compromised as all the RERs do not maintain long-term relationships. This conclusion is consistent with some previous research that focused on the existence of the G-PPP as a precondition for the creation of a monetary union (Sun and Simons (2011) ; Wilson and Choy (2007)). However, with respect to estimations results, two sub-groups seem to share a common long-run equilibrium suggesting that further monetary integration is possible. From this perspective, the Philippines has the highest integration because their RER is cointegrated with those of all other countries, and vice versa. Conversely, Singapore is the most isolated country in Asia, ceteris paribus, and it restricts the feasibility of a monetary union among the ASEAN-5 countries. More importantly, we find positive β coefficients in all cases, thus implying that countries that share a long-term equilibrium are symmetrically adjusted to macroeconomic disturbances. These long-run elasticities are particularly high between the RERs of Indonesia and other countries. For instance, the RER of Indonesia, based on the currency basket, has appreciated 1.31%, 1.21% and 1.62% following a 1% appreciation of the Malaysia ringgit, the Philippine peso and the Thai baht, respectively. Interestingly, we do not observe any substantial difference in the value of the β coefficient regardless of the value of the numeraire, which shows that the yen, like the dollar, is important for these countries. These results confirm the evidence from a number of previous studies that emphasize the need to promote a common currency basket with a large weight for the yen (Mishra and Sharma (2010) ; Ogawa and Kawasaki (2008)).

The value of the gamma coefficient that corresponds to the integration order of residuals yields information about the speed of adjustment towards the equilibrium. The greater is the difference between δ and γ, the greater is the reduction of the dimension, and the smaller are shocks between the RERs that lead to a persistent deviation from the long run equilibrium. Indeed, different types of cointegration relationships can be found depending on the responsiveness of equilibrium errors to the shocks (see section 3). From the value of the gamma coefficient, it can be seen that weak fractional cointegration relationships are observed in all cases. Clearly, deviations are higly persistent but there is an evident tendency of reversion toward equilibrium. Effectively, the long memory coefficient of the cointegrating errors range from 0.716 to 0.872 for basket-based RERs, from 0.717 to 0.865 for yen-based RERs and from 0.738 to 0.893 for USD-based RERs. Again, no obvious dissimilarities are observed according to the numeraire.

4.2. Comparative analysis

In a last step, we calculate estimates for the pre-crisis and post-crisis samples to take into account possible change in the relationship. Indeed, one major consequence of the crisis has been the emergence of a higher degree of policy cooperation in trade and monetary areas that result in greater interdependence of economic structures and contribute to the co-movements of real output variables in the long-run (Sato and Zhang (2006)). Therefore, it is important to consider how the Asian financial crisis has influenced the real macroeconomic factors that guide the RERs in East Asia. By comparing the number of pairs fractionally cointegrated according to the sub-samples, we observe 23 pairs of cointegrated countries before the crisis and 29 after the crisis. The difference is even greater when the numeraire is the dollar supporting the conclusion proffered by Choudhry (2005). More importantly, structural reforms seem to have led to increased co-movements of their RERS as equilibrium errors respond faster to shocks when comparing gamma coefficients between the two sub-samples. This implies that deviations toward the common equilibrium are less persistent, thus suggesting that the real income process underlying the evolution of the RERs has been more interrelated since the crisis. However, the evidence is not sufficient enough to assert that a block indexed to the dollar emerges in Asia, or that these countries have operated as an OCA since the crisis. The post-crisis results suggest a complex pattern of long-term correlations, as no clear group seems to stand out in this subsample. We do find, however, two pair of cointegrated RERs when we use the USD numeraire: Thailand-Indonesia and Singapore-Malaysia. Again, the RERs of all countries are fractionally cointegrated with the one of the Philippines. Furthermore, we find the following cointegrated pairs: Thailand depends on the RERs of Malaysia, Singapore depends on RERs of Malaysia, Malaysia depends on RERs of Indonesia. Finally, the fractional cointegration results with respect of basket-based and yen-based RERs are significant for nine pairs of countries, thus suggesting, once again, that the potentiality of a monetary union based on a currency basket composed of the yen and the dollar is more relevant.

5. Conclusion

This article analyzed the feasibility of a monetary union in the ASEAN-5 countries by applying the theory of G-PPP, which examines the existence of long-term co-movements between real exchange rates. The analysis covers a period from January 1975 to December 2011 and two sub-periods to account for the effects of the Asian financial crisis. A new estimator of fractional cointegration was introduced to test the presence of a long-run equilibrium between the real exchange rates of countries in this area. In contrast to all extant studies on the G-PPP, which focused on traditional cointegration, our model takes into account a wider range of dynamic reversion towards the long-run equilibrium. Doing so allows us to capture a larger number of long-term relationships and to analyze more precisely the degree of monetary integration of the countries in this area. Our analysis highlights two different sub-groups of countries through which the formation of an OCA is possible from the perspective of the regional stability of the the real exchange rates. Nevertheless, our results also show that this long-term stability is fragile and that monetary integration of these countries is still weak with respect to the speed of the absorption of the disturbances on real income. However, we find that the equilibrium errors respond more quickly to shocks in the post-crisis sample resulting in less persistent deviations towards the common equilibrium. This is an encouraging result when we consider the desirability of a currency union in the future. In this context the choice of the numeraire is important and the results show that an exchange rate policy based on a basket of currencies is probably more effective in ensuring the regional stability of the real exchange rates. Actually, most of the East Asian countries operate under managed float regimes for their nominal exchange rate by controlling their margins of fluctuations. However, there is not yet a formal or informal forms of exchange rates policy coordination that may lead to strengthen intra-regional exchange rate stability and which could then establish a wide zone of relative price stability. Our results support the case for further monetary integration in the ASEAN-5 countries by the development of reforms that aim to increase intra-regional trade and real linkages, such as informal forms of exchange rates policy coordination where the yen and the dollar play an essential role. As intra-regional trade and FDI flows continue to increase, the region will become more interdependent than it would if paving the way for a full-fledged monetary union.

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Table 2: Estimation of equation (1) for the endogenous variable: Indonesia

Indonesia		1975:01 - 2011:12 (n = 444)				1975:01 - 1997:05 (n = 269)				1998:06 - 2011:12 (n = 162)			
x_t		Mala.	Phil.	Sing.	Thai.	Mala.	Phil.	Sing.	Thai.	Mala.	Phil.	Sing.	Thai.
Test LV		***	**		**					***	***	***	***
USD	β	1.474	1.144	1.821	1.672	1.197	0.067	0.721	1.025	2.418	0.773	1.145	1.941
		(0.071)	(0.111)	(0.130)	(0.096)	(0.085)	(0.113)	(0.138)	(0.184)	(0.317)	(0.211)	(0.266)	(0.231)
	δ	0.959	0.976	0.926	0.932	0.906	0.918	0.897	0.985	0.876	0.997	0.950	0.985
		(0.071)	(0.068)	(0.072)	(0.074)	(0.092)	(0.086)	(0.093)	(0.090)	(0.124)	(0.106)	(0.120)	(0.120)
	γ	0.738	0.867	0.954	0.893	0.948	1.014	1.041	0.986	0.561	0.505	0.494	0.553
		(0.087)	(0.071)	(0.071)	(0.083)	(0.105)	(0.082)	(0.083)	(0.086)	(0.112)	(0.116)	(0.119)	(0.131)
Test LV		***	**		**					***	***	***	***
YEN	β	1.219	1.186	1.589	1.454	1.174	0.923	1.277	1.361	0.857	0.393	0.489	1.310
		(0.043)	(0.062)	(0.068)	(0.058)	(0.042)	(0.067)	(0.062)	(0.062)	(0.210)	(0.198)	(0.257)	(0.205)
	δ	0.970	0.929	0.981	0.955	1.015	0.953	1.014	1.035	0.880	0.873	0.832	0.878
		(0.070)	(0.076)	(0.070)	(0.073)	(0.086)	(0.095)	(0.093)	(0.086)	(0.138)	(0.122)	(0.135)	(0.128)
	γ	0.769	0.850	0.939	0.865	0.902	0.963	1.016	0.906	0.544	0.572	0.569	0.401
		(0.088)	(0.071)	(0.075)	(0.094)	(0.115)	(0.091)	(0.092)	(0.101)	(0.112)	(0.116)	(0.118)	(0.141)
Test LV		***	**		**					***	***	***	***
BSK	β	1.314	1.211	1.856	1.626	1.213	0.702	1.251	1.464	1.343	0.406	0.601	1.719
		(0.054)	(0.082)	(0.093)	(0.075)	(0.051)	(0.092)	(0.089)	(0.096)	(0.337)	(0.242)	(0.328)	(0.242)
	δ	0.962	0.928	0.963	0.943	0.988	0.914	0.969	1.027	0.930	0.942	0.920	0.950
		(0.070)	(0.074)	(0.073)	(0.072)	(0.085)	(0.099)	(0.090)	(0.083)	(0.130)	(0.114)	(0.114)	(0.121)
	γ	0.752	0.854	0.945	0.862	0.909	0.993	1.031	0.919	0.509	0.547	0.538	0.443
		(0.088)	(0.067)	(0.075)	(0.091)	(0.112)	(0.084)	(0.088)	(0.099)	(0.109)	(0.118)	(0.119)	(0.138)

Note: The lag structure has been selected using a likelihood ratio test. The test always concludes in favor of a $VARFI(1)$.
***. **. *. indicate the significance of the LV's test at thresholds of 1%. 5% and 10% respectively.
Results in parenthesis are standard errors.

Table 3: Estimation of equation (1) for the endogenous variable: Malaysia

Malaysia		1975:01 - 2011:12 (n = 444)				1975:01 - 1997:05 (n = 269)				1998:06 - 2011:12 (n = 162)			
x_t		Indo.	Phil.	Sing.	Thai.	Indo.	Phil.	Sing.	Thai.	Indo.	Phil.	Sing.	Thai.
Test LV		***	**		***	*	*				***	** *	***
USD	β	0.341	0.551	1.004	0.826	0.363	0.171	0.849	0.811	0.103	0.399	0.555	0.492
		(0.017)	(0.053)	(0.059)	(0.046)	(0.026)	(0.061)	(0.058)	(0.086)	(0.014)	(0.039)	(0.043)	(0.041)
	δ	0.937	0.976	0.926	0.932	1.018	0.918	0.897	0.985	0.569	0.997	0.950	0.985
		(0.069)	(0.068)	(0.078)	(0.072)	(0.083)	(0.093)	(0.093)	(0.088)	(0.122)	(0.112)	(0.120)	(0.103)
	γ	0.764	0.883	1.006	0.834	0.675	0.888	1.050	0.808	0.933	0.583	0.541	0.626
		(0.082)	(0.069)	(0.066)	(0.077)	(0.197)	(0.096)	(0.077)	(0.098)	(0.109)	(0.141)	(0.155)	(0.128)
Test LV		**	**		***		*					**	*
YEN	β	0.529	0.934	1.256	1.092	0.649	0.828	1.133	1.140	0.076	0.618	0.864	0.802
		(0.019)	(0.032)	(0.029)	(0.027)	(0.022)	(0.041)	(0.025)	(0.034)	(0.025)	(0.044)	(0.048)	(0.054)
	δ	0.958	0.929	0.981	0.955	1.065	0.953	1.014	-0.898	0.629	0.873	0.832	0.878
		(0.065)	(0.080)	(0.073)	(0.071)	(0.077)	(0.095)	(0.091)	(0.057)	(0.104)	(0.121)	(0.133)	(0.124)
	γ	0.717	0.863	1.034	0.774	0.763	0.901	1.078	0.744	0.874	0.811	0.665	0.780
		(0.104)	(0.069)	(0.065)	(0.081)	(0.145)	(0.089)	(0.079)	(0.112)	(0.145)	(0.124)	(0.138)	(0.120)
Test LV			**		***				***			**	*
BSK	β	0.437	0.824	1.289	1.044	0.536	0.638	1.123	1.157	0.059	0.415	0.606	0.529
		(0.018)	(0.042)	(0.041)	(0.037)	(0.024)	(0.054)	(0.040)	(0.049)	(0.017)	(0.043)	(0.058)	(0.053)
	δ	0.599	0.942	0.920	0.950	0.629	0.873	0.832	0.878	0.569	0.997	0.950	0.985
		(0.106)	(0.104)	(0.127)	(0.119)	(0.104)	(0.121)	(0.133)	(0.124)	(0.122)	(0.112)	(0.120)	(0.103)
	γ	0.717	0.872	1.025	0.792	0.710	0.918	1.073	0.758	0.950	0.766	0.707	0.743
		(0.097)	(0.063)	(0.060)	(0.079)	(0.138)	(0.099)	(0.074)	(0.109)	(0.128)	(0.128)	(0.144)	(0.132)

Note: The lag structure has been selected using a likelihood ratio test. The test always concludes in favor of a $VARFI(1)$.
***. **. *. indicate the significance of the LV's test at thresholds of 1%. 5% and 10% respectively.
Results in parenthesis are standard errors.

Table 4: Estimation of equation (1) for the endogenous variable: Phillipines

Phillipines		1975:01 - 2011:12 (n = 444)				1975:01 - 1997:05 (n = 269)				1998:06 - 2011:12 (n = 162)			
x_t		Indo.	Mala.	Sing.	Thai.	Indo.	Mala.	Sing.	Thai.	Indo.	Mala.	Sing.	Thai.
Test LV													
USD	β	0.172	0.358	0.603	0.490	0.025	0.176	0.437	0.231	0.067	0.937	0.882	0.806
		(0.017)	(0.035)	(0.051)	(0.046)	(0.036)	(0.063)	(0.080)	(0.097)	(0.024)	(0.098)	(0.061)	(0.067)
	δ	0.937	0.959	0.926	0.932	1.018	0.906	0.897	0.985	0.569	0.876	0.950	0.985
		(0.068)	(0.068)	(0.071)	(0.070)	(0.090)	(0.097)	(0.090)	(0.083)	(0.117)	(0.113)	(0.119)	(0.125)
	γ	0.919	0.920	0.906	0.889	0.910	0.902	0.865	0.894	0.987	0.835	0.888	0.756
		(0.072)	(0.068)	(0.073)	(0.073)	(0.088)	(0.086)	(0.090)	(0.086)	(0.109)	(0.107)	(0.123)	(0.126)
Test LV		***	**	**	***	**	*	**	*				**
YEN	β	0.385	0.701	1.021	0.879	0.460	0.737	0.955	0.957	0.036	0.900	1.057	1.007
		(0.020)	(0.024)	(0.030)	(0.028)	(0.032)	(0.036)	(0.040)	(0.044)	(0.030)	(0.062)	(0.059)	(0.050)
	δ	0.958	0.970	0.981	0.955	1.065	1.015	1.014	1.035	0.629	0.880	0.832	0.878
		(0.065)	(0.070)	(0.074)	(0.082)	(0.074)	(0.073)	(0.091)	(0.084)	(0.109)	(0.126)	(0.131)	(0.127)
	γ	0.775	0.819	0.828	0.753	0.714	0.794	0.791	0.796	0.865	0.824	0.851	0.643
		(0.093)	(0.076)	(0.080)	(0.093)	(0.133)	(0.113)	(0.103)	(0.131)	(0.121)	(0.101)	(0.135)	(0.125)
Test LV		**	**	**	***	*	*	**	*				*
BSK	β	0.276	0.568	0.925	0.747	0.288	0.571	0.834	0.841	0.021	0.846	0.984	0.897
		(0.018)	(0.029)	(0.045)	(0.038)	(0.034)	(0.044)	(0.062)	(0.063)	(0.026)	(0.089)	(0.078)	(0.078)
	δ	0.948	0.962	0.963	0.943	1.055	0.988	0.969	1.027	0.599	0.930	1.829	0.950
		(0.064)	(0.069)	(0.069)	(0.073)	(0.079)	(0.085)	(0.097)	(0.081)	(0.109)	(0.131)	(0.122)	(0.130)
	γ	0.809	0.841	0.835	0.785	0.751	0.814	0.804	0.793	0.938	0.840	0.860	0.687
		(0.084)	(0.073)	(0.078)	(0.089)	(0.123)	(0.106)	(0.095)	(0.128)	(0.115)	(0.108)	(0.129)	(0.134)

Note: The lag structure has been selected using a likelihood ratio test. The test always concludes in favor of a $VARFI(1)$.
***. **. *. indicate the significance of the LV's test at thresholds of 1%. 5% and 10% respectively.
Results in parenthesis are standard errors.

Table 5: Estimation of equation (1) for the endogenous variable: Singapore

Singapore		1975:01 - 2011:12 (n = 444)				1975:01 - 1997:05 (n = 269)				1998:06 - 2011:12 (n = 162)			
x_t		Indo.	Mala.	Phil.	Thai.	Indo.	Mala.	Phil.	Thai.	Ind.	Mal.	Phi.	Tha.
Test LV				**				**	**		***	*	
USD	β	0.173	0.416	0.385	0.488	0.138	0.539	0.271	0.661	0.075	0.925	0.625	0.730
		(0.012)	(0.023)	(0.032)	(0.031)	(0.026)	(0.039)	(0.045)	(0.068)	(0.020)	(0.077)	(0.042)	(0.047)
	δ	0.937	0.959	0.976	0.932	1.018	0.906	0.918	0.985	0.569	0.876	0.997	0.985
		(0.069)	(0.069)	(0.066)	(0.073)	(0.082)	(0.085)	(0.084)	(0.089)	(0.118)	(0.115)	(0.112)	(0.123)
	γ	0.947	0.979	0.803	0.860	0.936	1.046	0.828	0.820	0.930	0.659	0.736	0.707
		(0.077)	(0.066)	(0.082)	(0.081)	(0.098)	(0.082)	(0.102)	(0.103)	(0.121)	(0.140)	(0.160)	(0.144)
Test LV							*	*	**		***		**
YEN	β	0.355	0.652	0.704	0.772	0.481	0.770	0.728	0.932	0.032	0.758	0.642	0.782
		(0.015)	(0.015)	(0.021)	(0.021)	(0.023)	(0.019)	(0.030)	(0.027)	(0.024)	(0.043)	(0.034)	(0.046)
	δ	0.958	0.970	0.929	0.955	1.065	1.015	0.953	1.035	0.629	0.880	0.873	0.878
		(0.064)	(0.069)	(0.076)	(0.076)	(0.077)	(0.086)	(0.095)	(0.082)	(0.102)	(0.140)	(0.120)	(0.117)
	γ	0.957	1.045	0.884	0.884	0.930	1.084	0.864	0.800	0.812	0.617	0.779	0.618
		(0.087)	(0.067)	(0.083)	(0.091)	(0.110)	(0.084)	(0.101)	(0.110)	(0.141)	(0.127)	(0.160)	(0.181)
Test LV				**				**	**		***		**
BSK	β	0.255	0.537	0.558	0.636	0.333	0.674	0.558	0.864	0.022	0.671	0.533	0.652
		(0.013)	(0.017)	(0.027)	(0.026)	(0.025)	(0.024)	(0.038)	(0.040)	(0.018)	(0.061)	(0.036)	(0.052)
	δ	0.948	0.962	0.928	0.943	1.055	0.988	0.914	1.027	0.599	0.930	0.942	0.950
		(0.066)	(0.071)	(0.073)	(0.073)	(0.080)	(0.088)	(0.094)	(0.082)	(0.111)	(0.130)	(0.112)	(0.121)
	γ	0.964	1.028	0.855	0.879	0.914	1.069	0.855	0.792	0.908	0.708	0.738	0.594
		(0.090)	(0.066)	(0.084)	(0.068)	(0.113)	(0.081)	(0.099)	(0.110)	(0.120)	(0.123)	(0.170)	(0.183)

Note: The lag structure has been selected using a likelihood ratio test. The test always concludes in favor of a $VARFI(1)$.
***. **. *. indicate the significance of the LV's test at thresholds of 1%. 5% and 10% respectively.
Results in parenthesis are standard errors.

Table 6: Estimation of equation (1) for the endogenous variable: Thailand

Thailand		1975:01 - 2011:12 (n = 444)				1975:01 - 1997:05 (n = 269)				1998:06 - 2011:12 (n = 162)			
x_t		Indo.	Mala.	Phil.	Sing.	Indo.	Mala.	Phil.	Sing.	Indo.	Mala.	Phil.	Sing.
Test LV			**	*	*					***		**	
USD	β	0.245	0.524	0.476	0.749*	0.115	0.298	0.082	0.384	0.158	0.907	0.634	0.808
		(0.014)	(0.030)	(0.044)	(0.046)	(0.020)	(0.033)	(0.033)	(0.038)	(0.018)	(0.081)	(0.048)	(0.057)
	δ	0.937	0.959	0.976	0.926*	1.018	0.906	0.918	0.897	0.569	0.876	0.997	0.950
		(0.069)	(0.068)	(0.068)	(0.071)	(0.082)	(0.094)	(0.082)	(0.094)	(0.120)	(0.124)	(0.108)	(0.122)
	γ	0.898	0.829	0.827	0.883*	0.940	0.916	0.971	0.927	0.966	0.827	0.657	0.837
		(0.080)	(0.079)	(0.083)	(0.079)	(0.100)	(0.094)	(0.092)	(0.093)	(0.117)	(0.111)	(0.134)	(0.130)
Test LV			***	**	**		**		*				
YEN	β	0.418	0.724	0.781	0.988	0.479	0.724	0.684	0.877	0.120	0.701	0.607	0.786
		(0.016)	(0.018)	(0.027)	(0.028)	(0.021)	(0.021)	(0.031)	(0.024)	(0.022)	(0.054)	(0.038)	(0.043)
	δ	0.958	0.970	0.929	0.981	1.065	1.015	0.953	1.014	0.629	0.880	0.873	0.832
		(0.063)	(0.072)	(0.075)	(0.073)	(0.076)	(0.084)	(0.081)	(0.085)	(0.113)	(0.138)	(0.115)	(0.136)
	γ	0.830	0.771	0.816	0.863	0.826	0.813	0.969	0.878	0.829	0.835	0.740	0.837
		(0.110)	(0.087)	(0.083)	(0.082)	(0.116)	(0.107)	(0.090)	(0.092)	(0.138)	(0.106)	(0.125)	(0.126)
Test LV			***	*	**								
BSK	β	0.325	0.630	0.657	0.923	0.326	0.580	0.469	0.727	0.112	0.594	0.498	0.679
		(0.015)	(0.022)	(0.032)	(0.037)	(0.022)	(0.024)	(0.035)	(0.036)	(0.016)	(0.068)	(0.042)	(0.056)
	δ	0.948	0.962	0.928	0.963	1.055	0.988	0.914	0.969	0.599	0.930	0.942	0.920
		(0.065)	(0.069)	(0.074)	(0.072)	(0.075)	(0.082)	(0.093)	(0.092)	(0.115)	(0.134)	(0.114)	(0.124)
	γ	0.834	0.785	0.822	0.868	0.845	0.858	0.994	0.914	0.912	0.875	0.779	0.868
		(0.107)	(0.083)	(0.080)	(0.082)	(0.113)	(0.100)	(0.088)	(0.091)	(0.128)	(0.108)	(0.114)	(0.125)

Note: The lag structure has been selected using a likelihood ratio test. The test always concludes in favor of a $VARFI(1)$.
***. **. *. indicate the significance of the LV's test at thresholds of 1%. 5% and 10% respectively.
Results in parenthesis are standard errors.

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