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Abstract

This paper examines stock market integration between the ASEAN five and the US and China, respectively, over the period from November 2002 to March 2018. The linkages between both aggregate and financial sector stock indices (both weekly and monthly) are analysed using fractional integration and fractional cointegration methods. Further, recursive cointegration analysis is carried out for the weekly series to study the impact of the 2007-8 global financial crisis and the 2015 China stock market crash on the pattern of stock market co-movement. The main findings are the following. All stock indices exhibit long memory. There is cointegration between the ASEAN five and the US but almost none between the former and China, except between Indonesia and China in the case of the financial sector. The 2007-8 global financial crisis and the 2015 Chinese stock market plunge weakened the linkages between the ASEAN five and both China and the US. The implications of these results for market participants and policy makers are discussed.

JEL-Codes: C220, C320, G110, G150.

Keywords: Asian stock markets, financial integration, fractional integration, fractional cointegration.

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1. Introduction

In the most recent decades, globalisation has led to increasingly stronger linkages between stock markets and a higher degree of cross-country co-movement, with the US playing a dominant role. In the specific case of the Asian region it appears that China is also becoming an important driver of stock market movements (Shu et al., 2015). There have been a number of cases when developments in China have sent shock waves across Asia. For instance, the financial stimulus package announced by the Chinese authorities in November 2008 was followed on the same day by a 3.4% hike in Asian shares; between June and August 2015 the Chinese stock market plunged and sharp drops also occurred in other major Asian markets (e.g., -22% in Hong Kong, -11% in Japan, -7% in Korea, and -14% in Singapore).

It is often argued that while the US exerts its influence on the Asian financial markets through funding cost, portfolio rebalancing and risk appetite channels, China affects these economies mainly through trade linkages (Glick and Hutchison, 2013; Shu et al., 2015; Arslanalp et al., 2016). However, the 2015 turbulence across the Asian stock markets, seemingly originating from China, raised the question of whether financial linkages between these countries have now become more important (Rhee, 2015; IMF, 2016). Whilst the US impact on global (including Asian) financial markets is well documented, there are fewer studies examining financial integration between China and the other Asian economies. In particular, Shu et al. (2015, 2018) and Fang and Bessler (2018) found an increasingly significant impact of the Chinese stock market on others in the region.

Examining such linkages has important policy implications. Growth in Asia is increasingly driven by financial liberalisation rather than by exports (Arslanalp et al., 2016). China's impact on the Asian financial markets is likely to grow further with the ongoing internationalisation of the Renminbi and the removal of barriers to capital flows. Therefore, it is important for regional institutions such as the Association of Southeast Asian Nations

(ASEAN) and the Asian Development Bank to monitor the degree of co-movement between the stock markets of China and of the other Asian economies in order to design macro-prudential policies aimed at safeguarding financial stability in the presence of possible shocks originating from the larger economies in the region such as China. It is noteworthy that banks dominate in the financial sector in Asia, and intra-regional cross-border banking activity has in fact increased since the 2007-8 crisis global financial crisis (Remolona and Shim, 2015), following which the ASEAN countries adopted in 2014 the ASEAN Banking Integration Framework.

The present paper provides new evidence on financial linkages between China, the US and five ASEAN economies. Its contribution is threefold. First, it applies fractional integration and cointegration techniques to test for co-movement between stock markets in those countries. Such methods are more general than the standard ones based on the classical dichotomy between stationary I(0) and non-stationary I(1) series and have not been previously used in this context. Second, it assesses the impact of the 2007-8 global financial crisis as well as of the 2015 Chinese stock market plunge on financial integration between the same set of countries by carrying out recursive cointegration tests; these issues have also not been thoroughly investigated in earlier studies. Finally, in view of the fact that banks dominate Asia's financial systems, and given the intensified intra-regional cross-border banking activity since the 2007-8 crisis, this study also focuses specifically on the financial sector indices of stock markets. This topic has not been previously examined in depth, despite its direct relevance to policy makers given the regional banking integration framework recently adopted by the ASEAN members in the wake of the 2007-8 crisis.

The structure of the paper is as follows. Section 2 reviews the existing literature on Asian cross-market linkages and the role of China in the region. Section 3 outlines the empirical methods used for the analysis. Section 4 describes the data and discusses the main empirical results. Section 5 presents the results from the recursive cointegration analysis. Section 6 offers some concluding remarks.

2. Literature Review

The financial integration literature has mainly focused on the US as an international player. However, given China's growing size and increasingly important financial role, some recent studies have started analysing its regional influence. In particular, some of them have examined the currency markets and highlighted the rising impact of the Renminbi on the Asian currencies (e.g., Shu, 2010, Subramanian and Kessler, 2012; Fratzscher and Mehl, 2014; Shu et al., 2015; Caporale et al., 2018). Concerning the money market, Liu et al. (2013) found that real interest parity holds for ten East Asian countries that are significantly influenced by external factors originating from China. As for the bond markets, He et al. (2009) reported that Hong Kong's market is more integrated with the Chinese one during tranquil periods but more aligned with the US one in turbulent times; more recently, however, Shu et al., (2018) examined bond markets in China and eleven other Asian-Pacific economies and found that China remains a negligible player in the region.

As far as stock markets are concerned, some recent contributions have investigated cross-market linkages in Asia focusing in particular on the role of China (e.g., Lean and Smyth, 2014; Shu et al., 2015; Arslanalp et al., 2016; Singh and Kaur, 2016; Shu et al., 2018). Fan et al. (2009) and Srinivasan et al. (2013) analysed the linkage between China and other international markets. These studies use various methods such as cointegration (i.e., Srinivasan et al., 2013; Lean and Smyth, 2014; Singh and Kaur, 2016), structural vector autoregressions (SVARs) (Shu et al., 2015; Shu et al., 2018), Capital Asset Pricing Model (CAPM) based regression analysis (Arslanalp et al., 2016), and Markov-Switching Vector Error Correction Models (MS-VECM) (Fan et al., 2009). The overall conclusion is that

China's influence within and outside the Asian region has been steadily increasing in recent years but is not yet as prominent as that of the US.

Regarding the impact of the 2007-8 global financial crisis, Glick and Hutchison (2013) investigated stock market linkages between China and the other Asian countries during the period June 2005-October 2012 and found stronger linkages during the crisis. Further, in August 2015 a number of Asian markets slumped following the sharp drop in the Chinese one, but stock market co-movements between China and Asia during that period have hardly been analysed. To our knowledge, the only available evidence is due to Fang and Bessler (2018) and Ahmed and Huo (2018); the former estimated cointegration and linear non-Gaussian acyclic models (LiNGAM) and found that the Chinese market had a strong negative impact on the other Asian markets, and the latter employed Bayesian VAR and BEKK-GARCH specifications and found stronger volatility spillovers from China to most Asia-Pacific stock markets during the turbulent period.

None of the above studies employs fractional integration/cointegration methods, despite their ability to detect long-run linkages even in the presence of sizeable short-run deviations from equilibrium (Dolatabadi et al., 2015). Specifically, the equilibrium error could be mean-reverting without being exactly I(0), as a fractionally integrated error term will also display mean-reverting behaviour (Hosking, 1981). To our knowledge, there are only two papers estimating a fractional integration VECM model in the present context. Specifically, Chen et al. (2006) test for fractional cointegration between three pairs of stock markets (i.e., India-US, India-China and China-US), finding a dominant role for the US; Yi et al. (2009) show that the Chinese stock market has stronger ties with the Hong Kong one than with the US one. The present study is the first to employ fractional cointegration methods to examine the stock market linkages between China, five core ASEAN economies (i.e., Indonesia, Malaysia, Philippines, Singapore and Thailand), and the US.

Another noticeable gap in the literature is the lack of evidence on the impact of the 2007-8 global financial crisis and the 2015 Chinese stock market crash on financial linkages in the region (with the exception of the contributions by Glick and Hutchison (2013), Fang and Bessler (2018) and Ahmed and Huo (2018) already mentioned). Our recursive cointegration analysis sheds light on the dynamic pattern of stock market co-movements between China and the rest of Asia and between the US and Asia during tranquil and turbulent periods[1]. It also provides evidence on regional pull and global push factors as in Shu et al. (2018).

Finally, only limited attention has been paid in previous studies to intraregional banking in Asia. This has been on a steady rise in the past two decades, especially after the 2007-8 global financial crisis, when a prolonged period of low global long-term interest rates, new bank regulation and efforts to repair balance sheets opened up opportunities for banks in Asia to expand their activity within the region (Remolona and Shim, 2015). The recent push for regional bank integration by the ASEAN member countries is likely to lead to even greater intra-regional lending. Chinese banks have become an increasingly important provider of cross-border bank credit to borrowers within Asia. According to the BIS international banking statistics (2018), borrowers located in China accounted for almost half of the total increase in bank lending to emerging Asia in the first quarter of 2018. As pointed out by Koch and Remolona (2018), the common lender (i.e., China) channel generates more risk for emerging Asia if the shocks originate from the Asian borrowers themselves. Therefore a careful analysis, such as the one below, of the linkages between the largely bank-dominated financial sectors of China and the other Asian markets is most needed.

[1] To our knowledge the only other study using recursive cointegration for China is due to Yang (2003), who examined segmentation within China's six stock markets (i.e., Shanghai A-share, Shanghai B-share, Shenzheng A-share, Shenzheng B-share, Hong Kong H-share, red chip stock) and did not find long-run linkages between them.

3. Empirical Methodology

To understand the fractional integration and cointegration methods used below some definitions are useful. In particular, a covariance stationary process $\{x_t, t = 0, \pm 1, \ldots\}$ is defined to be I(0) or to exhibit short memory if the infinite sum of the autocovariances is finite, i.e., assuming $\gamma_u = E(x_t - Ex_t)(x_{t+u}-Ex_t))$, if

$$\sum_{u=-\infty}^{\infty} \left| \gamma_u \right| \quad < \quad \infty.^{2} \tag{1}$$

This category includes the stationary and invertible AutoRegressive Moving Average (ARMA) models. In this context, y_t is said to be integrated of order d or I(d) if d is the value of the differencing parameter required to make it stationary and the process can be expressed as:

$$(1 \; - \; L)^{d} \, y_t \quad = \quad x_t \, , \quad t \quad = \quad 0, \pm 1, \ldots, \tag{2}$$

where L is the lag operator and x_t is I(0). Then, if d > 0, y_t is said to exhibit long memory, so called because of the high degree of dependence between observations far apart in time. The polynomial on the left-hand side in (2) can be expressed in terms of its Binomial expansion, such that for all real d,

$$(1 - L)^d = \sum_{j=0}^{\infty} \binom{d}{j} (-1)^j L^j = 1 - dL + \frac{d(d-1)}{2} L^2 - \cdots,$$

and thus, equation (2) can be expressed as

$$y_t = d y_{t-1} - \frac{d(d-1)}{2} y_{t-2} + \cdots + x_t.$$

Thus, the higher the value of d, the higher is the level of dependence between the observations, and, if d is fractional, y_t depends on all its past history. These processes were original proposed in the 1980s by Granger (1980, 1981), Granger and Joyeux (1980) and Hosking (1981) and became very popular for the analysis of economic time series from the

[2] Alternatively, in the frequency domain, x_t is defined to be I(0) if its spectral density fucntion is positive and bounded at all frequencies.

late 1990s onwards (see, e.g., Baillie, 1996; Gil-Alana and Robinson, 1997; Mayoral, 2006; Gil-Alana and Moreno, 2012; Abbritti et al., 2016; etc.).

Below we estimate d using two approaches, parametric and semi-parametric respectively, both based on the Whittle function in the frequency domain (Dahlhaus, 1989). The former uses a testing procedure due to Robinson (1994) that is the most efficient method in the Pitman sense against local departures from the null. Another advantage of this approach is that it has a standard null limit distribution and this asymptotic behaviour holds regardless of the inclusion of deterministic terms such as intercepts or linear trends. Moreover, it is valid for any real value of d and therefore it is not necessary to differentiate the series in case of non-stationary behaviour. The latter follows Robinson (1995) and is based on a local Whittle function with frequencies degenerating to zero. This method requires stationarity and therefore the analysis will be conducted on the first-differenced series, adding 1 to the estimated value of d.[3]

Fractional cointegration is a natural extension of the concept of fractional cointegration to the multivariate case. In this paper we focus on bivariate relationships; following Engle and Granger (1987), two series, y_{1t} and y_{2t}, are said to be cointegrated if a) both of them are integrated of the same order, say d, and b) if there exists a linear combination of the two which is integrated of a smaller order, say d - b, with b < 0. Note that in the original paper by Engle and Granger (1987) the two parameters d and b were allowed to be fractional, though most of the empirical applications since then have focused on the integer case, with d = b = 1; only more recent studies have also allowed for fractional values. Below we test the order of integration of the differentials between the ASEAN stock indices and those of US and China, i.e. we impose the cointegrating vector to be exactly (1, -1) since

[3] Other semi-parametric methods (Abadir et al., 2007; Phillips and Shimotsu, 2004, 2005; Shao, 2010) require additional user-chosen parameters and they may be too sensitive to the choice of these numbers. Using some of these methods produced very similar results to those obtained with Robinson's (1995) method which we report.

estimating the cointegrating coefficients would require the computation of finite sample critical values for the confidence bands, which would be computationally very intensive.

4. Data and Empirical Results

The series used for the analysis are weekly aggregate stock market indices as well as financial sector indices for China, the ASEAN five (i.e., Indonesia, Malaysia, Philippines, Singapore and Thailand) and the US from November 2002 to March 2018. The data source is Datastream. Figure 1 displays the weekly series for the aggregates; all of them peak in 2007, decline sharply in 2008 and then increase steadily, though less obviously in the case of China, until mid-2014; another peak in 2015 is followed by a fall, more pronounced in the Asian countries than in the US, and then steady growth in 2016-2017 (with the exception of the Philippines) and a slight dip towards the end of the sample period. Financial sector indices (weekly) are plotted in Figure 2; they exhibit similar patterns to those in Figure 1, although the dips in 2007 and 2015 are more pronounced than for the aggregate indices. As a robustness check we also examine monthly data for both the aggregate and financial sector indices (Figures 3 and 4). These display similar patterns to the weekly ones plotted in Figure 1 and 2. All series are logged for the empirical analysis below.

4.1. Weekly Aggregate Stock Indices Evidence

We start by analysing the weekly series and estimating the following model:

$$y_t = \alpha + \beta t + x_t, \qquad (1 - L)^{d_o} x_t = u_t, \qquad t = 1, 2, ..., \tag{3}$$

under the two cases of uncorrelated and autocorrelated errors, in the latter case using a non-parametric approach due to Bloomfield (1973) that produces autocorrelations decaying at an exponential rate as in the ARMA case. Table 1a displays the Whittle estimates of d, along with the 95% confidence intervals corresponding to the non-rejection values of d using

Robinson's (1994) parametric approach, for the three standard cases of no deterministic terms ($\alpha = \beta = 0$ in (3)), an intercept (α estimated and $\beta = 0$), and an intercept and a linear time trend (i.e., α and β both unknown and estimated from the data).

Under the assumption of white noise errors there appears to be mean reversion ($d < 1$) in the case of the US; the unit root null hypothesis (i.e., $d = 1$) cannot be rejected for Indonesia, the Philippines and Thailand, whilst it is rejected in favour of higher orders of integration ($d > 1$) for China, Malaysia and Singapore. When allowing for autocorrelation as in Bloomfield (1973), evidence of I(d) behaviour with $d > 1$ is found for China, Indonesia, Singapore and Thailand, and evidence of I(1) behaviour for Malaysia, the Philippines and the US. On the whole, there is evidence of I(d) behaviour with $d = 1$ or $d > 1$ in all cases, which implies that shocks have permanent effects.

[Insert Tables 1a and 1b about here]

Table 1b displays the estimates of d from the semiparametric approach. The results are reported for a selected group of bandwidth parameters, from 25 to 31. The I(1) hypothesis cannot be rejected for any of them in the case of Indonesia, and for some in the case of Malaysia and the US. These results are consistent with the parametric ones, since they also provide evidence of I(d) behaviour with $d = 1$ or $d > 1$ in all cases.

Next we analyse the cointegrating relations with respect to both China and the US by considering the differentials between those two countries and the ASEAN ones. The results vis-à-vis China are reported in Table 2a. There is no evidence of mean reversion ($d < 1$), and thus of convergence, regardless of whether the residuals are assumed to be a white noise or an autocorrelated process, i.e. there is no evidence of convergence with respect to China.

[Insert Tables 2a and 2b about here]

Table 2b displays the results vis-à-vis the US. In this case, unlike the previous one, mean reversion and therefore convergence is found for all series. However, the adjustment process is slow, since the order of integration is large in all cases, ranging between 0.8 and 0.9. When allowing for autocorrelation, there is weaker evidence of convergence, and the unit root null cannot be rejected in some cases (i.e., Indonesia and Thailand)

In brief, our analysis provides some weak evidence of convergence vis-à-vis the US, but none vis-à-vis China. This is consistent with recent studies on Asian stock markets finding that global integration between the Asian and US stock markets dominates regional integration (Hinojales and Park, 2011; Park and Lee, 2011; Kim et al., 2011; Kim and Lee, 2012; Park, 2013).

4.2 Weekly Financial Sector Indices Evidence

Table 3a displays the Whittle estimates of d for the financial sector indices for the three standard cases of no deterministic terms ($\alpha = \beta = 0$ in (3)), an intercept (α estimated and $\beta = 0$), and an intercept and a linear time trend; d is estimated to be higher than 1 only in the case of Malaysia and Singapore, whilst in all other cases the unit root null cannot be rejected except for the US, where mean reversion ($d < 1$) is found. When allowing for autocorrelation as in Bloomfield (1973), a unit root is found in Indonesia, Thailand and the US, whilst in the other cases the order of integration is I(d, $d > 1$). Therefore, as in the case of the aggregate stock indices, there is evidence of shocks having permanent effects given the I(d) behaviour of the series with $d = 1$ or $d > 1$ in all cases.

[Insert Tables 3a and 3b about here]

The estimates of d with the semiparametric approach are displayed in Table 3b, again for the same bandwidth parameters (i.e., from 25 to 31). The I(1) hypothesis is not rejected for Indonesia and Thailand. In the other cases, the estimated values of d are above 1. Thus, there is evidence of I(d) behaviour with $d = 1$ or $d > 1$ in all cases, which is consistent with the parametric estimates in Table 3a.

The results of the fractional cointegrating analysis vis-à-vis China and the US are reported in Table 4a and 4b respectively. In the case of China, if no autocorrelation is allowed, there is no evidence of mean reversion (d < 1) and thus of convergence, except for some weak one in the case of Indonesia. When allowing for autocorrelation, no evidence of mean reversion is found in any single case.

[Insert Tables 4a and 4b about here]

By contrast, the results for the US (Table 4b) suggest mean reversion and thus convergence for all five ASEAN economies. The order of integration is relatively large in all cases (ranging between 0.8 and 0.9), which implies slow convergence. Under the assumption of autocorrelation, there is evidence of convergence only for Malaysia and Singapore while the unit root null cannot be rejected in the case of Indonesia, the Philippines and Thailand.

To sum up, as in the case of the aggregate stock indices, there is only weak evidence of convergence between the financial sector indices, with only one case of mean reversion vis-à-vis China (i.e., Indonesia with autocorrelated disturbances) compared to none when using the aggregate series. Further, the estimates of d for the financial sector indices vis-à-vis China (see Table 4a) are smaller than those for the aggregate indices (see Table 2a), which suggests closer integration of the financial sector.

This can be explained by Chinese banks becoming an increasingly important provider of cross-border credit to the Asian economies (Remolona and Shim, 2015; Koch and Remolona, 2018). As pointed out by McGuire and van Rixtel (2012), the bulk of the increase in international claims extended by outside area banks is to borrowers in China, and is largely consistent with the increase in the total assets of Chinese banks' foreign offices in Asia. Given the Chinese government's support for companies' "go global" policies, the accelerating internationalisation of the Renminbi, and new policy initiatives such as "One

Belt, One Road", the integration between the Chinese and ASEAN five's banking sector is likely to intensify further (Arslanalp et al., 2016).

4.3. Robustness Checks: Monthly Data

In this section we repeat the analysis using monthly data as a robustness check. [4] The fractional integration and cointegration results are presented in Tables 5a-8b.

The estimates based on the parametric method of Robinson (1994) are displayed in Table 5a for the aggregate indices and Table 7a for the financial sector indices. Under the white noise assumption, the unit root null hypothesis (i.e., d = 1) cannot be rejected for any of the aggregate indices but Singapore (d = 1.13), for which it is rejected in favour of d > 1. Concerning the financial indices (i.e., Table 7a), the unit root null is never rejected. When allowing for autocorrelated errors, the unit root null is not rejected in any single case. Therefore there is evidence of I(d) behaviour with d = 1 in practically all cases.

[Insert Tables 5a, 5b, 7a and 7b about here]

The semiparametric results are reported in Table 5b and 7b for a selected group of bandwidth parameters, from 11 to 17. They are consistent with the parametric ones, and provide further evidence of I(d) behaviour with d = 1 in the majority of cases.

Next we test again for convergence. In the case of China (in Table 6a and 8a), it can be seen that the time trend is only required vis-à-vis Indonesia with both white noise and autocorrelated errors and for both aggregates and financial indices. No evidence of mean reversion (d < 1) is found in any single case.

[Insert Tables 6a, 6b, 8a and 8b about here]

[4] As argued by Raj and Dhal (2008), although daily data capture the speedy transmission of information, since both short- and long-run dynamic linkages matter for market integration (Voronkova (2004), Hassan and Naka (1996)), less frequent (e.g., weekly as in the present study) stock returns are useful to avoid the problem of non-synchronous trading in some thinly traded stock markets (Cha and Oh , 2000).

As for the US results, evidence of mean reversion and hence convergence is found for Malaysia in the case of the aggregate index, and for Thailand with both aggregate (Table 6b) and financial indices (Table 8b). However, the values of d are relatively large, which implies a very slow process of convergence. With autocorrelated disturbances, the only case of mean reversion for the financial indices is Thailand.

On the whole, the monthly results are consistent with the weekly ones: although there are fewer instances of mean reversion, again there is stronger evidence of convergence between the ASEAN five and the US compared to China, both for the aggregate and the financial sector indices. Lower frequency data such as monthly ones might not detect transient responses to innovations that may last for a short period only (Brailsford, 1996; Elyasiani, et al., 1998; Andersen, et al. 2002; Palamalai et al., 2013). Therefore, we shall employ weekly data for the recursive cointegration analysis.

5. Recursive Cointegration Analysis

In this section we carry out recursive cointegration analysis using weekly data. In particular, we start with a sample of 114 observation, with data ending on 31/12/2004. Then, we add 20 observations (weeks) at a time recursively. Figures 5 and 6 display the estimates of d along with the 95% confidence bands, once more using Robinson's (1994) approach for the aggregate and financial sector indices respectively.

Concerning the aggregate results (Figure 5), for the US evidence of mean reversion ($d < 1$) is found in all cases for Singapore and the Philippines, and for Malaysia, Indonesia and Thailand except for some initial values. However, for China, the I(1) hypothesis cannot be rejected in any single case.

[Insert Figure 5 about here]

More specifically, in all the aggregate cases, the estimates of d increase quickly during the period 2007-2008 and then start decreasing both vis-à-vis the US and China. This might reflect the impact of the 2007-8 global financial crisis that sent shockwaves to stock markets across the world. The rising estimates of d indicate that at that stage stock market linkages became weaker, especially between the ASEAN five and China, with d moving from below to above 1 in some cases (e.g., Singapore, Philippines and Thailand). This is in contrast to previous studies where stronger ASEAN-US (e.g., Shu et al, 2018) and ASEAN-China stock market linkages (e.g., Glick and Hutchison, 2013) were found.

One possible explanation for the weaker financial linkages during this period is the overreaction hypothesis, which suggests that, after a series of bad news, investors become over-pessimistic about the future and send stock prices to unjustifiably low levels (Barberis et al., 1998). Because of the overreaction of the Asian stock prices to the financial crisis originating from the US, mean-reversion became slower. It is noteworthy that, whilst the crisis started as a financial and liquidity one in the US, the transmission mechanism to Asia was international trade (Raj and Roy, 2014), with households and firms putting off consumptions because of their fear of a further worsening in economic conditions which lead to a sharp decline in trade volumes. The growing values in d are probably a reflection of the different nature of the crisis and of its manifestations in the US and in Asia. In the case of the ASEAN-China linkages, different policy response may have also contributed to the lack of co-movement during the crisis. Specifically, countries with stronger fiscal position and low inflation such as China responded aggressively to the crisis with a sizeable fiscal stimulus package; instead in many other Asian countries, which were constrained by their fiscal position and/or inflation performance (e.g., Philippines, Malaysia), monetary policy responses may have been more muted (Bernanke, 2009).

[Insert Figure 6 about here]

Concerning the financial sector indices (Figure 6), similar patterns are found to those of the aggregate indices (Figure 5). Specifically, concerning convergence with respect to the US, the estimated values of d for Malaysia, Singapore and Thailand are significantly below 1, and the same is true for Indonesia and the Philippines apart from some initial values. The impact of the 2007-8 crisis seems to have been less pronounced in the financial sector since the estimate of d increases less than in the case of the aggregate indices. This reflects the fact that the transmission of the crisis to Asia took place mainly through trade as opposed to financial channels (Essers, 2013). Regarding the ASEAN five-China linkages, the estimates of d for the financial sector indices (right panel of Figure 6) are lower than those for the aggregate indices (right panel of Figure 5) for all five countries. This indicates that although there is no evidence of mean reversion for either the aggregate or financial sector indices, the financial sector is much closer to becoming integrated than the aggregate one. This can probably be attributed to the rise of regional banking in Asia and to the leading role played by Chinese banks as the dominant regional credit provider (Remolona and Shim, 2015).

Further, in the second half of the sample period, in the case of China (i.e., right panels of Figures 5 and 6) the estimates of d remain stable until 2015 and then start decreasing and stabilising for both the aggregate and financial sector indices. As previously mentioned, in 2015 there was a stock market crash in China and turbulence in several Asian markets in the region shortly after. This confirms the influence of the Chinese stock market on others in the region (Shu et al., 2015; Fang and Bessler, 2018). However, the rising estimates of d during the Chinese stock market crisis period suggesting weaker linkages between financial sectors. This is in contrast to the findings of Fang and Bessler (2018) and Ahmed and Huo (2018), who reported stronger volatility spillovers between China and the other Asian countries during the 2015 crisis period. Weaker linkages might be attributed to the overreaction effect mentioned above and also to the fact that the main cause of the stock market crash in China was the sharp increase in the number of ordinary Chinese people investing in it: according to Reuters, more than 40 million new stock accounts were opened between June 2014 and May 2015, with many of the account holders making highly leveraged purchases using borrowed money. The Chinese government also adopted a series of unconventional measures to prop up the market such as allowing companies voluntarily to suspend trading of their shares to prevent a further loss of value. At the height of the crisis in July 2015, more than 50 percent of all companies listed on the Shanghai and Shenzhen stock exchanges stopped trading (Salidjanova, 2015). Caporale et al (2016) found a similar rising pattern of the estimates of d during the volatile crisis period in their study of European and US stock market integration. Moreover, Figures 5 and 6 show that the 2015 stock market turmoil in China does not seem to have affected the ASEAN five-US financial linkages and has had a much less noticeable impact on the Asian economies compared to the 2007-8 global financial crisis, which suggests that China does not yet have a significant influence on global markets despite its growing regional power (as found by Shu et al., 2018).

Finally, the estimates of d in the case of the ASEAN five-US linkages have been consistently rising after the 2007-8 crisis (albeit gradually) until the end of the sample period (left panels of Figures 5 and 6). A similar trend is not present in the case of the ASEAN five-China linkages (i.e., right panel of Figures 5 and 6). This indicates co-movement between the stock markets of the ASEAN five and the US has been declining over time, consistently with the findings of Glick and Hutchison (2013) and Wu et al. (2015), who reported that the transmission of US equity returns to Asian countries has been decreasing since the global financial crisis.

6. Conclusions

This paper examines stock market integration between the ASEAN five and the US and China, respectively, over the period from November 2002 to March 2018. The linkages between both aggregate and financial sector stock indices (both weekly and monthly) are analysed using fractional integration and fractional cointegration methods. Further, recursive cointegration analysis is carried out for the weekly series to study the impact of the 2007-8 global financial crisis and the 2015 China stock market crash on the pattern of stock market co-movement. Our results can be summarised as follows.

First, the fractional integration tests suggest the presence of long-memory properties in all stock indices. This result is robust to using both parametric and semiparametric methods. Second, the cointegrating regression results provide evidence of cointegration between the ASEAN five and the US but almost none between the former and China, except between Indonesia and China in the case of the financial sector. The latter evidence, combined with the estimates of d between the ASEAN five and China being lower for all financial sector indices compared to the aggregate ones, implies that there is closer integration between the ASEAN five and China in the bank-dominated financial sectors than in the case of the aggregate stock markets. Third, the recursive cointegration analysis shows a significant impact of the 2007-8 global financial crisis and the 2015 Chinese stock market plunge on the ASEAN five-US and ASEAN five- China stock market integration. In contrast to most previous studies, we find that weaker stock market linkages resulted from these two events, possibly because of the overreaction of stock prices, the different nature and causes of two crises, and the different policy responses. Fourth, the 2015 China stock market turbulence had a much less significant impact on the Asian economies and did not affect much the financial linkages between the ASEAN five and the US. Finally, the increasing recursive estimates of d between the ASEAN five and the US, for both the aggregate and the financial sector indices, indicate that stock market linkages have become weaker over time. Similar results are found at a monthly frequency.

Our findings have several implications. First, the limited evidence of cointegration between China and the ASEAN five stock markets suggests that there is an opportunity for Chinese/Asian investors to diversify their portfolio by investing in the Chinese/Asian stock markets. Moreover, diversification is more effective across the aggregate stock markets than their financial sectors as the latter are more closely integrated, especially in the case of China and Indonesia. On the other hand, given the high degree of integration between the US and the Asian stock markets, both at the aggregate and financial sector level, diversification would not be effective in this case.

Second, since Asian stock markets appear to be more integrated with the US than with China, more regional agreements in addition to the recent regional initiatives (e.g., the Chiang Mai Initiative in 2000, the Asian Bond Market Initiative (ABMI) in 2003, the new ABMI roadmap in 2008, the Chiang Mai Initiative Multilateralization in 2012) are desirable to promote further financial cooperation in the region. For instance, from the point of view of financial regulation, Ananchotikul et al. (2015) point out that regulatory differences (e.g., in investor protection, bankruptcies procedures) within Asia may hinder regional financial integration and therefore more harmonisation should be pursued in this area to promote further integration.

Third, since both the 2007-8 global financial crisis and the 2015 China stock market crash had a negative influence on the financial integration process of the ASEAN stock markets with the Chinese and the US ones, appropriate measures should be taken such as building a greater financial resource pool (drawing on Asia’s rich foreign exchange reserve for instance) and strengthening the surveillance system in order to construct an effective regional facility that would provide emergency in the case of future financial crises (Asian Development Bank, 2012).

Finally, the rising of China as a regional economic power has played an important role in linking stock markets across the region (Glick and Hutchison, 2013). This phenomenon is particularly important in view of the declining degree of integration between the ASEAN five and the US found in our study. The recent emergence of China as the common lender in the bank-dominated financial sector in Asia is a possible consequence of the new regional banking integration framework, which is likely to be strengthened by Chinese policies such as “go global” and “One Belt, One Road”. At the same time, as pointed out by Remolona and Shim (2015), monetary authorities in Asia need to have macro-prudential policies in place to deal with the risks of financial instability originating from the Chinese financial sector, as in the case of the 2015 stock market crash.

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Table 1a. Estimates of d for each series without and with autocorrelation: weekly aggregate stock indices

	no autocorrelation			with autocorrelation		
Countries	No regressors	An intercept	A linear time trend	No regressors	An intercept	A linear time trend
CHINA	0.99 (0.95, 1.04)	**1.07 (1.03, 1.12)**	1.07 (1.03, 1.12)	0.98 (0.92, 1.07)	**1.12 (1.04, 1.19)**	1.12 (1.04, 1.19)
INDONESIA	0.99 (0.95, 1.04)	0.99 (0.95, 1.03)	**0.99 (0.95, 1.03)**	0.99 (0.92, 1.08)	1.10 (1.01, 1.19)	**1.10 (1.01, 1.19)**
MALAYSIA	0.99 (0.95, 1.05)	1.04 (1.00, 1.09)	**1.04 (1.00, 1.09)**	0.98 (0.92, 1.08)	1.06 (0.99, 1.16)	**1.06 (0.99, 1.16)**
PHILIPPINES	0.99 (0.95, 1.04)	0.99 (0.95, 1.04)	**0.99 (0.95, 1.04)**	0.98 (0.91, 1.07)	1.07 (0.99, 1.17)	**1.06 (0.99, 1.17)**
SINGAPORE	0.99 (0.95, 1.05)	**1.07 (1.03, 1.12)**	1.07 (1.03, 1.12)	0.98 (0.91, 1.07)	**1.12 (1.04, 1.19)**	1.11 (1.04, 1.19)
THAILAND	0.99 (0.95, 1.04)	1.02 (0.97, 1.06)	**1.01 (0.97, 1.06)**	0.99 (0.92, 1.08)	**1.11 (1.02, 1.19)**	1.10 (1.02, 1.19)
U.S.	0.99 (0.95, 1.05)	0.94 (0.90, 0.98)	**0.94 (0.90, 0.98)**	0.98 (0.92, 1.07)	0.99 (0.93, 1.08)	**0.99 (0.93, 1.08)**

Note: We report the estimates of d in the model given by equation (3). In brackets are the 95% bands for the non-rejection values of d. Values in bold indicate the most significant model for each series according to the deterministic terms and the type of I(0) disturbance.

Table 1b. Estimates of d for each series using a semiparametric Whittle method: weekly aggregate stock indices

	25	26	27	28	29	30	31
CHINA	**1.281***	**1.289***	**1.285***	**1.325***	**1.303***	**1.307***	**1.331***
INDONESIA	1.067	1.049	1.024	1.037	1.063	1.079	1.107
MALAYSIA	1.140	1.115	1.128	1.136	1.146	**1.174***	**1.194***
PHILIPPINES	1.131	1.132	1.148	**1.163***	1.143	**1.180***	**1.217***
SINGAPORE	1.246	**1.249***	**1.193***	**1.227***	**1.243***	**1.268***	**1.286***
THAILAND	**1.170***	**1.194***	1.154	**1.160***	**1.169***	**1.177***	**1.205***
U.S.	1.104	1.103	1.094	1.124	**1.156***	**1.184***	**1.198***
Lower 5%	0.835	0.838	0.841	0.844	0.847	0.849	0.852
Upper 5%	1.164	1.161	1.158	1.155	1.152	1.150	1.147

Note: * and in bold: Evidence of I(d) with d > 1 at the 5% level.

Table 2a. Estimates of d from the fractional cointegration regressions vis-à-vis China: weekly aggregate stock indices

	no autocorrelation			with autocorrelation		
Countries	No regressors	An intercept	A linear time trend	No regressors	An intercept	A linear time trend
INDONESIA	0.98 (0.94, 1.04)	0.98 (0.94, 1.02)	**0.98 (0.94, 1.02)**	0.99 (0.92, 1.07)	**1.06 (0.98, 1.14)**	1.06 (0.98, 1.14)
MALAYSIA	1.02 (0.98, 1.06)	**1.03 (0.99, 1.07)**	1.03 (0.99, 1.07)	1.06 (0.99, 1.15)	**1.08 (1.00, 1.15)**	1.08 (1.00, 1.15)
PHILIPPINES	1.00 (0.96, 1.05)	**1.00 (0.96, 1.05)**	1.00 (0.96, 1.05)	1.05 (0.98, 1.15)	**1.06 (0.99, 1.16)**	1.06 (0.99, 1.16)
SINGAPORE	1.01 (0.96, 1.06)	**1.01 (0.96, 1.06)**	1.01 (0.96, 1.06)	1.05 (0.97, 1.13)	**1.04 (0.97, 1.13)**	1.04 (0.97, 1.13)
THAILAND	1.02 (0.98, 1.07)	**1.02 (0.98, 1.08)**	1.02 (0.98, 1.08)	1.08 (1.01, 1.19)	**1.08 (1.01, 1.17)**	1.08 (1.01, 1.17)

Note: In brackets are the 95% band for the non-rejection values of d. Values in bold indicate the most significant model for each series according to the deterministic terms and the type of I(0) disturbance. Values in bold and red indicate the rejection of the unit root null.

Table 2b. Estimates of d from the fractional cointegration regressions vis-à-vis the US: weekly aggregate stock indices

	no autocorrelation			with autocorrelation		
Countries	No regressors	An intercept	A linear time trend	No regressors	An intercept	A linear time trend
INDONESIA	0.99 (0.94, 1.04)	0.88 (0.84, 0.92)	**0.89 (0.83, 0.92)**	0.97 (0.90, 1.06)	0.96 (0.89, 1.02)	**0.97 (0.90, 1.02)**
MALAYSIA	0.96 (0.91, 1.01)	**0.84 (0.80, 0.87)**	0.84 (0.80, 0.87)	0.97 (0.91, 1.06)	**0.88 (0.82, 0.94)**	0.88 (0.82, 0.94)
PHILIPPINES	0.98 (0.94, 1.03)	0.87 (0.84, 0.92)	**0.87 (0.84, 0.92)**	0.99 (0.93, 1.08)	0.92 (0.86, 0.99)	**0.92 (0.86, 0.99)**
SINGAPORE	0.98 (0.94, 1.03)	**0.86 (0.82, 0.90)**	0.86 (0.82, 0.90)	0.99 (0.92, 1.08)	**0.94 (0.88, 0.99)**	0.94 (0.88, 1.01)
THAILAND	0.97 (0.92, 1.02)	0.88 (0.83, 0.92)	**0.88 (0.83, 0.92)**	0.96 (0.89, 1.06)	**0.96 (0.88, 1.07)**	0.96 (0.88, 1.07)

Note: In brackets the 95% band for the non-rejection values of d. Values in bold indicate the most significant model for each series according to the deterministic terms and the type of I(0) disturbance. Values in bold and red indicate the rejection of the unit root null.

Table 3a. Estimates of d for each series without and with autocorrelation: weekly financial sector indices

	no autocorrelation			with autocorrelation		
Countries	No regressors	An intercept	A linear time trend	No regressors	An intercept	A linear time trend
CHINA	0.99 (0.94, 1.04)	**1.02 (0.98, 1.07)**	1.02 (0.98, 1.07)	0.97 (0.91, 1.06)	**1.09 (1.01, 1.18)**	1.09 (1.01, 1.18)
INDONESIA	0.99 (0.94, 1.04)	0.95 (0.90, 1.00)	**0.95 (0.90, 1.00)**	0.99 (0.91, 1.07)	1.01 (0.92, 1.14)	**1.01 (0.92, 1.13)**
MALAYSIA	0.99 (0.95, 1.04)	1.06 (1.02, 1.11)	**1.06 (1.02, 1.11)**	0.98 (0.91, 1.07)	1.09 (1.00, 1.17)	**1.09 (1.00, 1.17)**
PHILIPPINES	0.99 (0.94, 1.05)	1.02 (0.98, 1.07)	**1.02 (0.98, 1.07)**	0.98 (0.90, 1.05)	**1.08 (1.00, 1.17)**	1.08 (1.00, 1.17)
SINGAPORE	0.99 (0.95, 1.05)	**1.07 (1.02, 1.12)**	1.07 (1.02, 1.12)	0.98 (0.91, 1.07)	**1.12 (1.04, 1.21)**	1.12 (1.04, 1.21)
THAILAND	0.99 (0.94, 1.04)	0.98 (0.94, 1.04)	**0.98 (0.94, 1.04)**	0.98 (0.91, 1.07)	1.01 (0.94, 1.11)	**1.01 (0.94, 1.11)**
U.S.	0.99 (0.95, 1.05)	**0.94 (0.90, 0.98)**	0.94 (0.90, 0.98)	0.98 (0.92, 1.07)	**1.04 (0.96, 1.11)**	1.04 (0.96, 1.11)

Note: The same as for Table 1a.

Table 3b. Estimates of d for each series using a semiparametric Whittle method: weekly financial sector indices

	25	26	27	28	29	30	31
CHINA	1.135	**1.163***	**1.169***	**1.198***	**1.214***	**1.215***	**1.227***
INDONESIA	0.830	0.843	0.815	0.835	0.861	0.885	0.898
MALAYSIA	**1.178***	**1.174***	**1.192***	**1.196***	**1.201***	**1.229***	**1.243***
PHILIPPINES	1.073	1.105	1.078	1.113	1.130	1.146	**1.166***
SINGAPORE	1.140	**1.189***	**1.220***	**1.215***	**1.155***	**1.191***	**1.223***
THAILAND	1.039	1.048	1.079	1.100	1.132	1.140	1.107
U.S.	**1.173***	**1.197***	**1.224***	**1.232***	**1.259***	**1.265***	**1.282***
Lower 5%	0.835	0.838	0.841	0.844	0.847	0.849	0.852
Upper 5%	1.164	1.161	1.158	1.155	1.152	1.150	1.147

Note: The same as for Table 1b.

Table 4a. Estimates of d from the fractional cointegration regressions vis-à-vis China: weekly financial sector indices

	no autocorrelation			with autocorrelation		
Countries	No regressors	An intercept	A linear time trend	No regressors	An intercept	A linear time trend
INDONESIA	0.98 (0.93, 1.03)	0.95 (0.90, 0.99)	**0.95 (0.90, 0.99)**	0.96 (0.89, 1.05)	**0.99 (0.92, 1.09)**	0.99 (0.91, 1.09)
MALAYSIA	1.00 (0.97, 1.06)	**1.00 (0.96, 1.05)**	1.00 (0.96, 1.05)	1.06 (0.99, 1.15)	**1.05 (0.97, 1.12)**	1.05 (0.97, 1.14)
PHILIPPINES	0.98 (0.94, 1.03)	**0.99 (0.94, 1.03)**	0.99 (0.94, 1.03)	1.00 (0.92, 1.10)	**1.01 (0.94, 1.12)**	1.01 (0.94, 1.12)
SINGAPORE	0.97 (0.93, 1.02)	**0.97 (0.93, 1.02)**	0.97 (0.93, 1.02)	1.05 (0.96, 1.14)	**1.04 (0.96, 1.14)**	1.04 (0.96, 1.14)
THAILAND	1.01 (0.96, 1.06)	**1.01 (0.96, 1.06)**	1.01 (0.96, 1.06)	1.00 (0.93, 1.07)	**1.02 (0.95, 1.11)**	1.02 (0.95, 1.11)

Note: The same as for Table 2a.

Table 4b. Estimates of d from the fractional cointegration regressions vis-à-vis the US: weekly financial sector indices

	no autocorrelation			with autocorrelation		
Countries	No regressors	An intercept	A linear time trend	No regressors	An intercept	A linear time trend
INDONESIA	0.99 (0.95, 1.04)	0.87 (0.84, 0.91)	**0.87 (0.84, 0.91)**	0.99 (0.92, 1.08)	0.94 (0.88, 1.02)	**0.94 (0.89, 1.02)**
MALAYSIA	0.96 (0.92, 1.01)	0.87 (0.84, 0.91)	**0.87 (0.84, 0.91)**	0.99 (0.93, 1.08)	**0.94 (0.89, 0.99)**	0.94 (0.89, 1.00)
PHILIPPINES	0.99 (0.95, 1.04)	0.89 (0.85, 0.93)	**0.89 (0.85, 0.93)**	1.02 (0.94, 1.09)	0.94 (0.87, 1.02)	**0.94 (0.87, 1.02)**
SINGAPORE	0.98 (0.94, 1.03)	0.86 (0.80, 0.89)	**0.86 (0.80, 0.89)**	0.99 (0.95, 1.08)	**0.90 (0.86, 0.97)**	0.90 (0.86, 0.97)
THAILAND	0.99 (0.95, 1.04)	0.87 (0.83, 0.92)	**0.87 (0.83, 0.92)**	0.97 (0.93, 1.07)	**0.92 (0.86, 1.02)**	0.92 (0.86, 1.02)

Note: The same as for Table 2b.

Table 5a. Estimates of d for each series without and with autocorrelation: monthly aggregate stock indices

	no autocorrelation			with autocorrelation		
Countries	No regressors	An intercept	A linear time trend	No regressors	An intercept	A linear time trend
CHINA	0.97 (0.99, 1.09)	**1.08 (0.99, 1.19)**	1.08 (0.99, 1.19)	0.97 (0.82, 1.19)	**1.10 (0.92, 1.35)**	1.10 (0.92, 1.35)
INDONESIA	0.99 (0.89, 1.11)	1.06 (0.96, 1.19)	**1.06 (0.96, 1.18)**	0.97 (0.83, 1.17)	1.00 (0.70, 1.27)	**1.01 (0.84, 1.24)**
MALAYSIA	0.98 (0.89, 1.09)	1.03 (0.93, 1.14)	**1.03 (0.94, 1.14)**	0.97 (0.81, 1.16)	1.08 (0.85, 1.32)	**1.07 (0.88, 1.30)**
PHILIPPINES	0.97 (0.89, 1.09)	1.01 (0.92, 1.12)	**1.01 (0.92, 1.12)**	0.94 (0.80, 1.16)	1.15 (0.93, 1.39)	**1.13 (0.94, 1.36)**
SINGAPORE	0.97 (0.88, 1.09)	**1.13 (1.03, 1.26)**	1.13 (1.03, 1.26)	0.96 (0.81, 1.18)	**1.07 (0.81, 1.35)**	1.06 (0.86, 1.34)
THAILAND	0.99 (0.90, 1.10)	0.99 (0.89, 1.10)	**0.99 (0.89, 1.10)**	0.96 (0.83, 1.17)	1.08 (0.83, 1.37)	**1.09 (0.86, 1.35)**
U.S.	0.98 (0.89, 1.10)	0.98 (0.89, 1.10)	**0.98 (0.89, 1.10)**	0.96 (0.81, 1.14)	0.95 (0.80, 1.16)	**0.95 (0.80, 1.16)**

Note: The same as for Table 1a.

Table 5b. Estimates of d for each series using a semiparametric Whittle method: monthly aggregate stock indices

	12	13	14	15	16	17	18
CHINA	0.779	0.901	1.017	1.082	1.166	**1.238***	**1.299***
INDONESIA	0.869	0.857	0.900	0.946	0.975	0.994	1.024
MALAYSIA	0.826	0.889	0.970	1.033	1.105	**1.202***	**1.274***
PHILIPPINES	1.062	0.924	0.973	1.027	1.048	1.085	1.087
SINGAPORE	0.819	0.992	0.830	0.897	0.893	0.952	0.951
THAILAND	0.847	0.800	0.847	0.906	0.920	0.935	0.979
U.S.	0.940	0.881	0.932	0.986	1.022	1.036	0.997
Lower 5%	0.762	0.771	0.780	0.787	0.794	0.800	0.806
Upper 5%	1.237	1.228	1.219	1.212	1.205	1.199	1.193

Note: The same as for Table 1b.

Table 6a. Estimates of d from the fractional cointegration regressions vis-à-vis China: monthly aggregate stock indices

	no autocorrelation			with autocorrelation		
Countries	No regressors	An intercept	A linear time trend	No regressors	An intercept	A linear time trend
INDONESIA	0.92 (0.84, 1.03)	0.94 (0.85, 1.05)	**0.94 (0.86, 1.05)**	0.95 (0.77, 1.14)	0.92 (0.74, 1.11)	**0.93 (0.78, 1.12)**
MALAYSIA	1.00 (0.91, 1.12)	**0.97 (0.89, 1.08)**	0.97 (0.89, 1.08)	0.98 (0.82, 1.20)	**1.01 (0.86, 1.23)**	1.01 (0.86, 1.23)
PHILIPPINES	0.97 (0.88, 1.08)	**0.95 (0.87, 1.06)**	0.95 (0.87, 1.06)	0.96 (0.81, 1.14)	**0.99 (0.85, 1.16)**	0.99 (0.84, 1.16)
SINGAPORE	0.97 (0.88, 1.08)	**0.96 (0.88, 1.07)**	0.96 (0.88, 1.07)	0.97 (0.81, 1.17)	**1.01 (0.84, 1.22)**	1.01 (0.84, 1.22)
THAILAND	0.97 (0.91, 1.10)	**0.97 (0.89, 1.06)**	0.97 (0.89, 1.06)	1.04 (0.90, 1.22)	**1.07 (0.91, 1.24)**	1.07 (0.91, 1.23)

Note: The same as for Table 2a.

Table 6b. Estimates of d from the fractional cointegration regressions vis-à-vis the US: monthly aggregate stock indices

	no autocorrelation			with autocorrelation		
Countries	No regressors	An intercept	A linear time trend	No regressors	An intercept	A linear time trend
INDONESIA	0.97 (0.88, 1.08)	0.96 (0.89, 1.07)	**0.97 (0.90, 1.07)**	0.93 (0.80, 1.13)	1.03 (0.90, 1.24)	**1.04 (0.92, 1.21)**
MALAYSIA	0.98 (0.89, 1.09)	**0.88 (0.82, 0.97)**	0.88 (0.82, 0.97)	0.91 (0.79, 1.06)	**0.92 (0.82, 1.04)**	0.92 (0.82, 1.04)
PHILIPPINES	0.99 (0.90, 1.11)	0.92 (0.84, 1.02)	**0.92 (0.85, 1.02)**	0.95 (0.82, 1.14)	0.92 (0.80, 1.09)	**0.92 (0.81, 1.08)**
SINGAPORE	1.02 (0.92, 1.14)	**1.01 (0.93, 1.10)**	1.01 (0.93, 1.10)	0.96 (0.83, 1.15)	**1.08 (0.97, 1.28)**	1.08 (0.97, 1.29)
THAILAND	0.92 (0.83, 1.04)	0.82 (0.74, 0.93)	**0.83 (0.76, 0.93)**	0.90 (0.74, 1.08)	**0.94 (0.76, 1.15)**	0.95 (0.81, 1.15)

Note: The same as for Table 2b.

Table 7a. Estimates of d for each series without and with autocorrelation: monthly financial sector indices

	no autocorrelation			with autocorrelation		
Countries	No regressors	An intercept	A linear time trend	No regressors	An intercept	A linear time trend
CHINA	0.96 (0.83, 1.18)	**0.98 (0.80, 1.19)**	0.98 (0.80, 1.19)	0.96 (0.83, 1.18)	**0.98 (0.80, 1.19)**	0.98 (0.80, 1.19)
INDONESIA	0.97 (0.81, 1.15)	0.70 (0.62, 1.01)	**0.78 (0.62, 1.00)**	0.97 (0.81, 1.15)	0.70 (0.62, 1.01)	**0.78 (0.62, 1.00)**
MALAYSIA	0.97 (0.81, 1.16)	1.11 (0.91, 1.37)	**1.12 (0.93, 1.36)**	0.97 (0.81, 1.16)	1.11 (0.91, 1.37)	**1.12 (0.93, 1.36)**
PHILIPPINES	0.95 (0.81, 1.14)	1.02 (0.82, 1.30)	**1.02 (0.83, 1.30)**	0.95 (0.81, 1.14)	1.02 (0.82, 1.30)	**1.02 (0.83, 1.30)**
SINGAPORE	0.96 (0.82, 1.15)	1.00 (0.78, 1.31)	**1.00 (0.80, 1.30)**	0.96 (0.82, 1.15)	1.00 (0.78, 1.31)	**1.00 (0.80, 1.30)**
THAILAND	0.96 (0.82, 1.19)	0.98 (0.78, 1.30)	**0.99 (0.78, 1.30)**	0.96 (0.82, 1.19)	0.98 (0.78, 1.30)	**0.99 (0.78, 1.30)**
U.S.	0.97 (0.82, 1.14)	**0.96 (0.83, 1.14)**	0.96 (0.83, 1.14)	0.97 (0.82, 1.14)	**0.96 (0.83, 1.14)**	0.96 (0.83, 1.14)

Note: The same as for Table 1a.

Table 7b. Estimates of d for each series using a semiparametric Whittle method: monthly financial sector indices

	11	12	13	14	15	16	17
CHINA	0.921	1.021	1.120	1.172	1.197	1.125	1.140
INDONESIA	0.809	0.773	0.821	0.886	0.885	0.810	0.829
MALAYSIA	0.957	0.975	1.066	1.105	1.114	1.192	1.244
PHILIPPINES	1.038	0.809	0.862	0.924	0.943	0.959	0.945
SINGAPORE	0.790	0.741	0.797	0.864	0.842	0.897	0.920
THAILAND	0.670	0.656	0.686	0.742	0.786	0.810	0.864
U.S.	1.071	0.993	1.036	1.085	1.102	1.073	1.035
Lower 5%	0.762	0.771	0.780	0.787	0.794	0.800	0.806
Upper 5%	1.237	1.228	1.219	1.212	1.205	1.199	1.193

Note: The same as for Table 1b.

Table 8a. Estimates of d from the fractional cointegration regressions vis-à-vis China: monthly financial sector indices

	no autocorrelation			with autocorrelation		
Countries	No regressors	An intercept	A linear time trend	No regressors	An intercept	A linear time trend
INDONESIA	0.93 (0.85, 1.04)	0.92 (0.83, 1.03)	**0.92 (0.84, 1.03)**	0.96 (0.81, 1.15)	0.89 (0.75, 1.07)	**0.89 (0.76, 1.07)**
MALAYSIA	1.03 (0.94, 1.15)	**1.00 (0.91, 1.11)**	1.00 (0.91, 1.11)	0.96 (0.82, 1.15)	**0.98 (0.84, 1.17)**	0.98 (0.84, 1.17)
PHILIPPINES	0.91 (0.83, 1.01)	**0.92 (0.83, 1.03)**	0.92 (0.83, 1.03)	0.94 (0.81, 1.14)	**0.89 (0.76, 1.06)**	0.88 (0.75, 1.06)
SINGAPORE	0.93 (0.85, 1.05)	**0.93 (0.84, 1.04)**	0.93 (0.84, 1.04)	0.91 (0.76, 1.11)	**0.92 (0.75, 1.11)**	0.92 (0.76, 1.11)
THAILAND	0.91 (0.84, 1.01)	**0.95 (0.87, 1.05)**	0.95 (0.87, 1.05)	0.99 (0.85, 1.19)	**0.99 (0.84, 1.15)**	0.99 (0.84, 1.15)

Note: The same as for Table 2a.

Table 8b. Estimates of d from the fractional cointegration regressions vis-à-vis the US: monthly financial sector indices

	no autocorrelation			with autocorrelation		
Countries	No regressors	An intercept	A linear time trend	No regressors	An intercept	A linear time trend
INDONESIA	0.98 (0.89, 1.10)	0.93 (0.85, 1.03)	**0.93 (0.86, 1.03)**	0.95 (0.82, 1.14)	0.96 (0.85, 1.14)	**0.96 (0.85, 1.13)**
MALAYSIA	0.99 (0.89, 1.11)	**0.95 (0.88, 1.04)**	0.95 (0.88, 1.04)	0.94 (0.84, 1.10)	**0.96 (0.86, 1.08)**	0.96 (0.86, 1.08)
PHILIPPINES	1.00 (0.91, 1.12)	0.93 (0.84, 1.06)	**0.93 (0.84, 1.06)**	0.96 (0.83, 1.16)	0.86 (0.74, 1.05)	**0.86 (0.73, 1.05)**
SINGAPORE	1.00 (0.91, 1.12)	**0.94 (0.87, 1.03)**	0.94 (0.87, 1.04)	0.97 (0.84, 1.17)	**0.96 (0.86, 1.10)**	0.96 (0.86, 1.10)
THAILAND	0.97 (0.89, 1.09)	0.85 (0.78, 0.95)	**0.85 (0.78, 0.95)**	0.94 (0.82, 1.13)	0.82 (0.74, 0.96)	**0.82 (0.73, 0.96)**

Note: The same as for Table 2b.

Figure 1. Weekly aggregate stock indices plots: China, ASEAN five and US

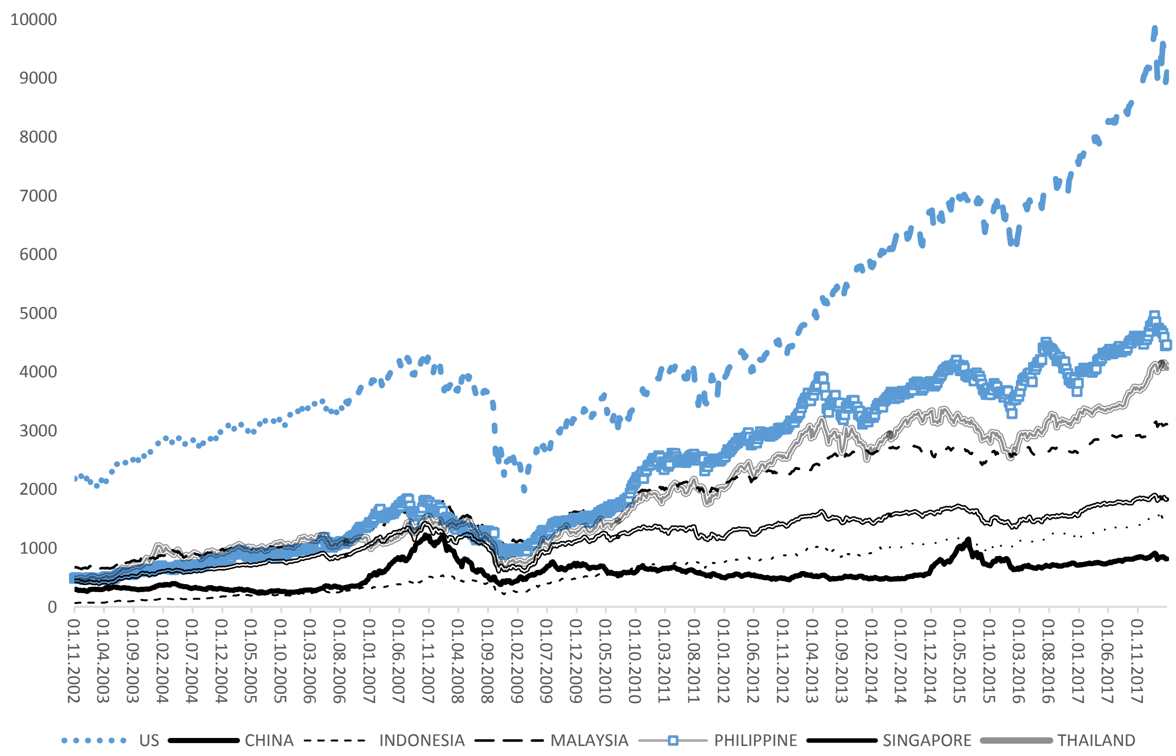


Figure 2. Weekly financial sector stock indices plots: China, ASEAN five and US

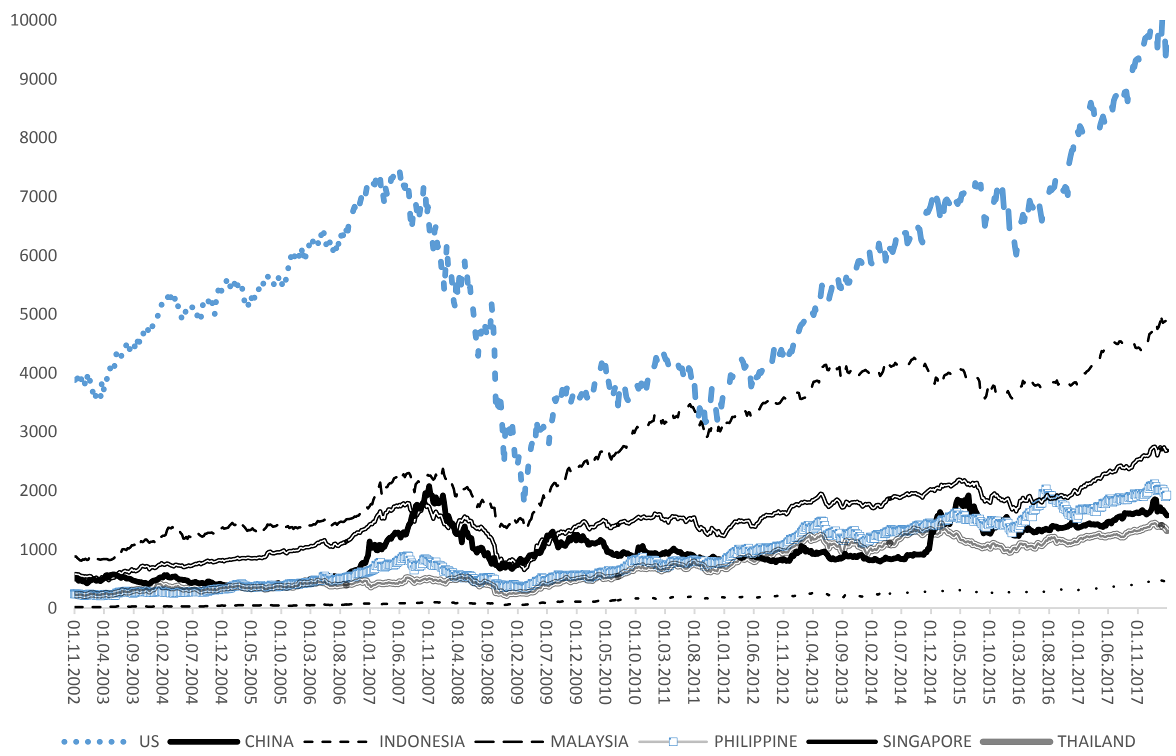


Figure 3. Monthly aggregate stock indices plots: China, ASEAN five and US

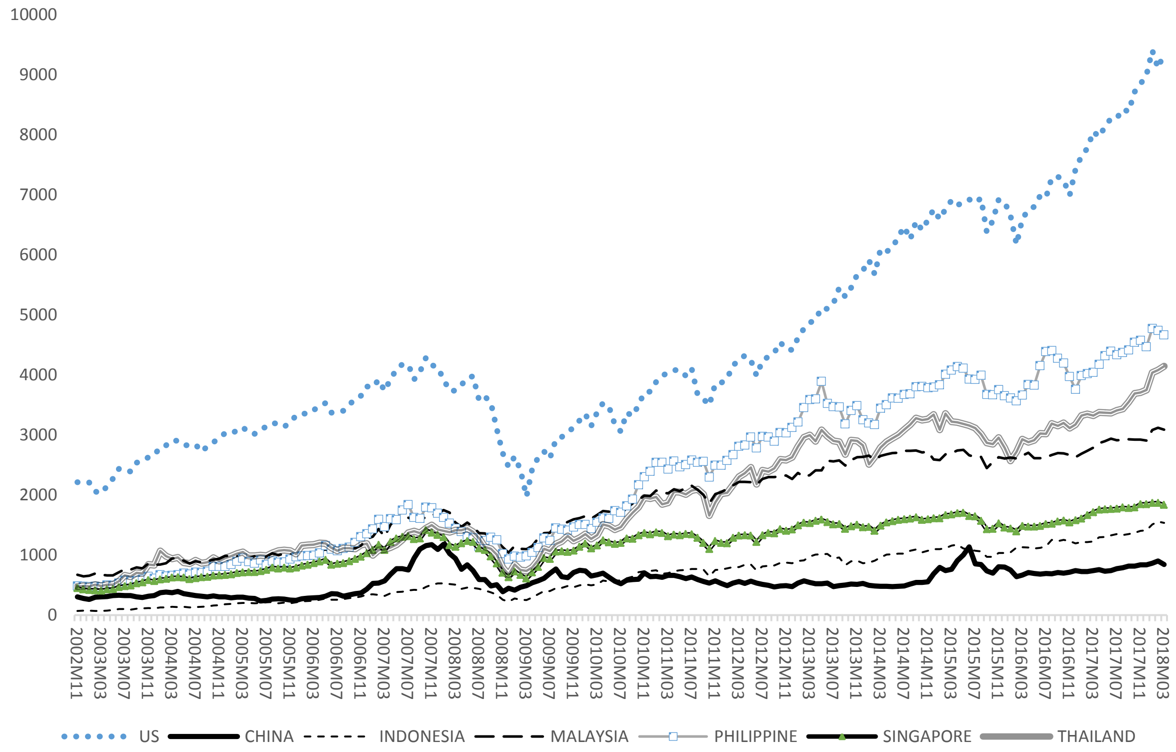


Figure 4. Monthly financial sector stock indices plots: China, ASEAN five and US

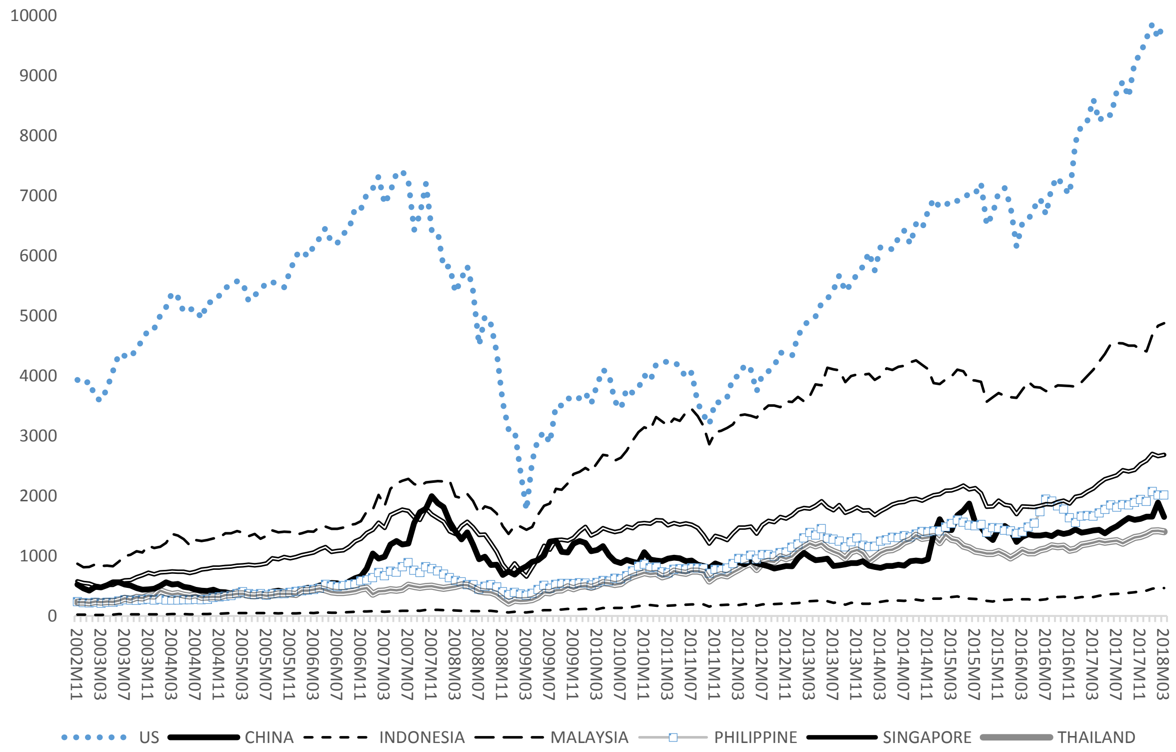
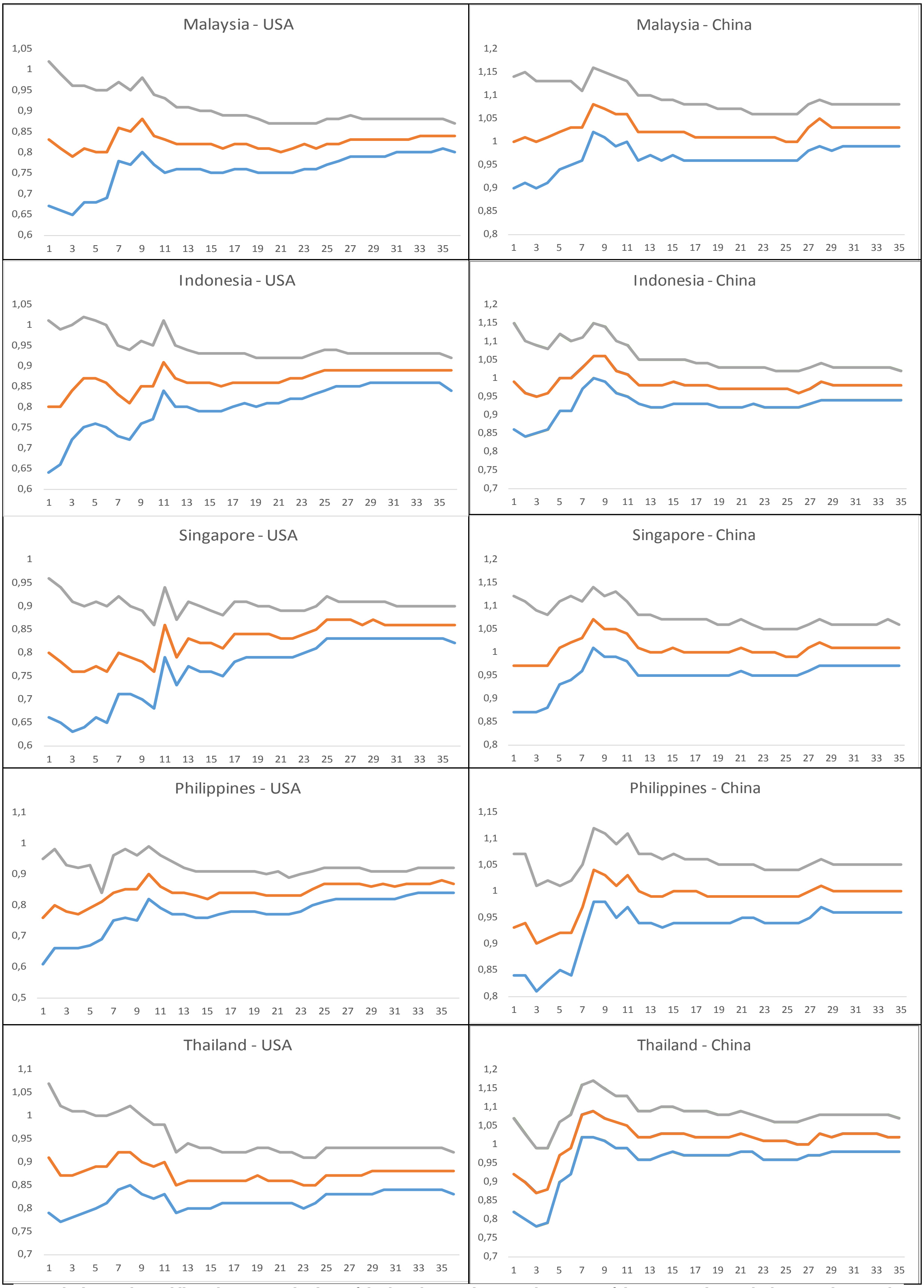


Figure 5. Recursive estimates of d from the fractional cointegration regressions: Aggregate indices



Note: The line in the middle is the estimated values of d. The other two lines are the 95% confidence intervals. On the horizontal axis is the number of rolling periods (each period is 20 weeks) used for each estimation. On the vertical are the estimated values of d.

Figure 6. Recursive estimates of d from the fractional cointegration regressions: Financial sector indices

Note: The same as for Figure 2.