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Stock Prices, Exchange Rates and Portfolio Equity Flows: A Toda-Yamamoto Panel Causality Test

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Structured Abstract

Purpose
The purpose of this paper is to develop a new framework to test the hypothesis that portfolio model predicts a negative correlation between stock prices and exchange rates in a trivariate transmission channel for foreign portfolio equity investment.

Design/methodology/approach
This paper utilizes panel data for eight economies to extend the Dumitrescu and Hurlin (2012) Granger non-causality test of heterogeneous panels to a trivariate model by integrating the Toda and Yamamoto (1995) approach to Granger causality.

Findings
The evidence suggests that stock prices Granger cause exchange rates and portfolio equity flows Granger cause exchange rates. However, the overall panel evidence casts doubt on the explicit trivariate model of portfolio balance model. The study shows that Indonesia may be the only case where stock prices affect exchange rates through portfolio equity flows.

Research limitations/implications
The proposed test does not account for potential asymmetries or structural shifts associated with the crisis period. To isolate the impact of the Asian Financial crisis, this paper rather splits the sample period in two sub-periods: pre- and post-crises. The sample period and countries are also limited due to the use of the balance of payment statistics.

Practical implications
The study casts doubt on the maintained hypothesis of a trivariate transmission channel, as posited by the portfolio model. Policy makers of an economy may integrate capital market and fiscal policies in order to maintain stable exchange rate.

Originality/value
This paper integrates a portfolio equity inflow variable into a single framework with stock price and exchange rate variables. It extends the Dumitrescu and Hurlin (2012)'s bivariate stationary Granger non-causality test in heterogeneous panels to a trivariate setting in the framework of Toda and Yamamoto (1995).

Keywords: stock prices; exchange rates; portfolio equity; Granger causality; heterogeneous panels

JEL Classification: F31, G14, G15

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STOCK PRICES, EXCHANGE RATES AND PORTFOLIO EQUITY FLOWS:
A Toda-Yamamoto Panel Causality Test

1. Introduction

A robust dynamic relationship between stock prices and exchange rates has been observed in Europe (Hau & Rey 2005) and Asia (Moore & Wang 2014). The phenomenon has attracted attention in the aftermath of the 1997 Asian financial crisis (Granger *et al.* 2000) and again has drawn a lot of interest from both academics and practitioners since the recent global financial crisis, as in Moore and Wang (2014), Inci and Lee (2014), Yang *et al.* (2014), Caporale *et al.* (2014), Groenewold and Paterson (2013), Liang *et al.* (2013), and Tsagkanos and Siriopoulos (2013). The liberalization of global financial asset transactions seems to be responsible for the dynamic relationship and has then led to the increased exposure of stock prices to exchange rate risks. Singh (1997) contends that an increase in stock market liquidity leads to more volatile foreign exchange markets. This linkage therefore has important implications for international portfolio management and the impact of stock markets on firm performance. Shocks, like exchange rate movements, may impact on equity markets and vice versa. From a policy perspective, it is important to identify causal effects between the monetary sector and the real economy. Stock and foreign exchange stock markets can impact on investment and GDP growth and causation may run both ways. Thus fiscal and monetary policies will be better informed by accounting for such potential causal links.

There are two main competing models that explain the relationship between stock prices and exchange rates; namely, the traditional approach models (Dornbusch & Fischer 1980) and the portfolio balance approach models (Frankel 1983).[1] According to the traditional models, exchange rates are determined by trade flows whereas the portfolio balance models posit that they are driven by financial market equilibrium conditions. Assuming a home country bias and imperfect substitute between domestic and foreign financial assets, Frankel (1983) argues that investors

[1] The former is also well known as flow-approach models; while the latter is also well known as stock-approach models or portfolio approach models. See Bahmani-Oskooee and Saha (2015) for the recent literature review on stock prices and exchange rates.

rebalance their portfolios according to the expected returns of both assets expressed in their domestic denominated currency. Under a floating exchange rate regime, an increase (decrease) in domestic asset prices will lead to an increase (decrease) in asset demand which then attracts capital inflows (outflows) and subsequently leads to an appreciation (depreciation) of the domestic currency.[2] Therefore, there is a negative unidirectional causality relationship from stock prices to exchange rates.[3] In contrast, the traditional approach postulates that the relationship may be positive or negative and that the direction of causality may start from stock prices to exchange rates or the other way around. An appreciation or depreciation of the exchange rate will affect both multinational firms (directly) and domestic firms (indirectly). Depending on whether a firm's main business is export or import-orientated, a change in the firm's performance due to the change in exchange rates leads to a change in investor valuation of the firm's stock price.

In the context of current integrated financial markets, the portfolio balance approach models seem to receive more empirical support and attention than its competitor (see for example Moore and Wang (2014), Caporale *et al.* (2014), Tsagkanos and Siriopoulos (2013), Filipe (2012), Lee *et al.* (2011), and Hau and Rey (2005)). However, these empirical studies mainly focus on a causal relationship without considering the impact of portfolio equity flows. In other words, they use a bivariate setting, not a trivariate setting. Portfolio equity flows are the neglected essential variable in portfolio balance models, and therefore these are susceptible to an omitted variable bias (Granger (1969) and Caporale *et al.* (2004)). Hau and Rey (2005) have developed a new approach to risk rebalancing associated with portfolio equity flows. They contend that portfolio flows are a key determinant of exchange rates and are induced by the need for rebalancing of the equity portfolio. Their model conjectures that (1) stock prices and exchange rates are negatively correlated and (2) a domestic currency depreciation and portfolio equity inflow is positively correlated. However, their empirical regression models only examine the impact of stock prices on exchange rates without incorporating a mediating role for equity portfolio inflows.

[2] A longer transmission channel is started from stock prices then to domestic investor wealth, money demands, interest rates, foreign capital flows, and finally exchange rates.

[3] An exchange rate is a value of one currency relative to another. The US dollar is commonly used as the base currency (the denominator). An appreciation (depreciation) in a domestic currency (the nominator) will decrease (increase) the exchange rate.

The portfolio flows are analysed separately from exchange rates and stock prices. This bivariate approach has also been used in other studies. Granger *et al.* (2000), Tsagkanos and Siriopoulos (2013) and Caporale *et al.* (2014) mention portfolio equity flows in interpretations of their results, but do not incorporate this variable into their empirical models. Granger *et al.* (2000), for instance, speculates that there is a capital expatriation from European equity markets to both the Gold market and the Asian equity markets. Meanwhile Caporale *et al.* (2014) deducts graphically that portfolio flows may be responsible in explaining their empirical findings that support the portfolio approach models.

One plausible reason for the neglect of portfolio equity flows may be due to non-availability of adequate data. Hau and Rey (2005) utilize the TIC data of Board of Governors of the Federal Reserve System, but this data only represent U.S. portfolio holdings of foreign securities. The net portfolio equity inflow data of the World Development Indicators seems to be a good alternative, but their annual nature makes it hard to have sufficiently long time series. This study uses quarterly data, i.e. the net portfolio investment of equity of the International Financial Statistics (IFS), to examine time-series properties of cross-country data and integrates the portfolio flow variable with the exchange rate and stock price variables. The expected relationship is depicted in Figure 1 and can be summarized as follows. Financial liberalisation enables investors to invest their money in any country and also withdraw the money from that country and move them to another country at any time without any restriction. A positive trend in stock prices in an economy will attract the global investors to enter that market. The activity of foreign investors (buying or selling) in the domestic equity market is reflected in the flows of portfolio equity in the balance of payment. The foreign equity flows then will affect exchange rates. Under the frameworks of Frankel (1983) and Hau and Rey (2005), this paper summarizes that a decrease (increase) in stock prices will lead to foreign equity capital outflows (inflows) and then eventually lead to depreciation (appreciation) of a domestic currency. In case of Granger causality, stock prices affect exchange rates through portfolio equity flows, i.e. stock prices affect portfolio equity flows and portfolio equity flows in turn affect exchange rates.

INSERT FIGURE 1 ABOUT HERE

In light of the above discussion, this paper aims to integrate portfolio equity inflows into a single framework with stock prices and the exchange rate. A better understanding of such trivariate links ought to contribute to better decisions by policy makers and investors. Policy makers in particular may integrate capital market and fiscal policies in order to maintain a stable exchange rate. This paper uses a trivariate Granger causality test to examine the relationship among the variables of interest. To this end, this paper extends the stationary bivariate non-causality test for heterogeneous panels of Dumitrescu and Hurlin (2012) to a trivariate setting with possible non-stationary variables. In particular we adapt the Toda and Yamamoto (1995) approach that allows non-stationary variables in a modified Granger causality test.

The contributions of this paper are (1) integrating a portfolio equity inflow variable into a single framework with stock price and exchange rate variables, (2) examining this in a panel setting which has better power (Carrion-i-Silvestre *et al.* 2005), and (3) extending the Dumitrescu and Hurlin (2012)'s bivariate stationary Granger non-causality test in heterogeneous panels to a trivariate setting in the framework of Toda and Yamamoto (1995). To the best of our knowledge, the only study that use a trivariate setting in the similar topic is that of Groenewold and Paterson (2013) which use commodity prices as the mediating variable in a time-series study for Australia.

The rest of the paper is organized as follows. Section 2 sets out the methodology. Section 3 describes the data. Section 4 provides the results. Finally section 5 concludes.

2. Methodology

Figure 1 shows that if net portfolio inflows (*EqFlows*) are omitted then a causality test between exchange rates (*Currency*) and stock prices (*Index*) may be invalid. A valid transmission follows the solid line rather than the dotted line. Therefore, a portfolio inflow variable must be included into a single framework with stock price and exchange rate variables.

A general dynamic interaction between stock prices, portfolio equity flows and exchange rates for each individual country i (i = 1,…,N) at time t (t = 1,…,T) can be modelled using three K-th order trivariate panel vector autoregressive (VAR) equations as follows:

$$Index_{i,t} = \alpha_{1i} + \sum_{p=1}^{K} \beta_{1i,p}\, Index_{i,t-p} + \sum_{p=1}^{K} \gamma_{1i,p}\, EqFlows_{i,t-p} + \sum_{p=1}^{K} \delta_{1i,p}\, Currency_{i,t-p} + \varepsilon_{1i,t} \quad (1)$$

$$EqFlows_{i,t} = \alpha_{2i} + \sum_{p=1}^{K} \beta_{2i,p}\, Index_{i,t-p} + \sum_{p=1}^{K} \gamma_{2i,p}\, EqFlows_{i,t-p} + \sum_{p=1}^{K} \delta_{2i,p}\, Currency_{i,t-p} + \varepsilon_{2i,t} \quad (2)$$

$$Currency_{it} = \alpha_{3i} + \sum_{p=1}^{K} \beta_{3i,p}\, Index_{i,t-p} + \sum_{p=1}^{K} \gamma_{3i,p}\, EqFlows_{i,t-p} + \sum_{p=1}^{K} \delta_{3i,p}\, Currency_{i,t-p} + \varepsilon_{3i,t} \quad (3)$$

where $\varepsilon_{1i,t}$, $\varepsilon_{2i,t}$, and $\varepsilon_{3i,t}$ denote individual white-noise errors and are assumed to be independently and normally distributed with $E(\varepsilon_{li,t}) = 0$ and $E(\varepsilon_{li,t}^2) = \sigma_{li}^2$, $\forall l = 1,2,3$. The errors are also independently distributed across countries where $E(\varepsilon_{li,t}\varepsilon_{lj,s}) = 0, \forall i \neq j, \forall t, s$. It is assumed that the models are heterogeneous panel data in which (1) α_{1i}, α_{2i} and , α_{3i} are fixed across time, (2) the lag order K, where K>0, is constant across equations, and (3) $\beta_{li,p}$, $\gamma_{li,p}$,and $\delta_{li,p}$, $\forall l = 1,2,3$ may vary either in an equation or across equations. This paper is interested at testing Granger causality between two variables of interest while controlling for the other variable.

In such panel VAR, there are at least three different estimation techniques that can be employed: a generalized method of moment (GMM) estimator (Holtz-Eakin *et al.* 1988; Love & Zicchino 2006), a seemingly unrelated regression (SUR) estimator (Konya 2006), and a multivariate least square estimator (Dumitrescu & Hurlin 2012).[4] Except for Love and Zicchino (2006) that is interested at the impulse-response function, Konya (2006) and Dumitrescu and Hurlin (2012) propose a different approach to Granger causality or non-causality test. Dumitrescu and Hurlin (2012) offer a bivariate non-causality test for heterogeneous panels which allows all coefficients to be different across cross-sections. On the other hand, Konya (2006)

[4] See Canova and Ciccarelli (2013) for a recent survey of the panel VAR literature.

offers a causality test for each individual country even though the estimation is conducted using a panel data setting. The test of Konya (2006) can be expanded to a trivariate setting but they treat the third variable as an auxiliary variable, not as an endogenous one such as in Love and Zicchino (2006).

This paper extends the test by Dumitrescu and Hurlin (2012) to a trivariate setting and relaxes the requirements that the variables of interest must be stationary variables or have the same order of integration. In a bivariate setting, Toda and Yamamoto (1995) argue that if one or both variables are non-stationary, a standard Granger causality test such as in Dumitrescu and Hurlin (2012) is not valid because the Wald test statistic does not follow its usual asymptotic chi-square distribution under the null hypothesis. To overcome this issue, they offer a different approach by introducing *m* additional lags to the time-series VAR (*K*) to ensure that the asymptotic distribution of the Wald test statistic still holds. However the extra *m* lags, which are the maximum order of integration of the time series variables, are not included in the Wald test. Cointegration tests are therefore needed for verification but they do not affect the Toda-Yamamoto test. With this simple alternative approach, a modified contrast matrix will have the same rank as the original one. This leads to the fact that the important properties of the panel Wald test statistic proposed by Dumitrescu and Hurlin (2012) still hold. A bivariate Toda Yamamoto approach in heterogeneous panels has also been offered by Emirmahmutoglu and Kose (2011). In contrast to Dumitrescu and Hurlin (2012), they use the Fisher test statistic which sums all individual country *p*-values to test the null hypothesis of Granger non-causality. The Fisher test statistic is claimed to have a chi-square distribution with 2*N* degrees of freedom when *N* is fixed and *T* reaches to infinity.

In a bivariate setting with both variables Y_i and X_{1i} being stationary, a general *K*-th order panel VAR equation can be written as: [5]

$$y_{i,t} = \alpha_i + \sum_{k=1}^{K} \beta_i^{(k)} y_{i,t-k} + \sum_{k=1}^{K} \gamma_i^{(k)} x_{1i,t-k} + \varepsilon_{i,t} \tag{4}$$

[5] This study here follows the notations used by Dumitrescu and Hurlin (2012).

with $K \in \mathbb{N}^*$, $\beta_i = \left(\beta_i^{(1)}, \dots, \beta_i^{(K)}\right)'$ and $\gamma_i = \left(\gamma_i^{(1)}, \dots, \gamma_i^{(K)}\right)'$. The coefficients of $\alpha_i, , \beta_i^{(k)}, \gamma_i^{(k)}$ are assumed to be constant in the time dimension. Individual residuals, $\varepsilon_{i,t}$, $\forall t = 1, \dots, T$, are assumed to be independently and normally distributed with $E(\varepsilon_{i,t}) = 0$ and $E(\varepsilon_{i,t}^2) = \sigma_i^2$. The residuals are also independently distributed across groups, $\varepsilon_i = (\varepsilon_{i,1}, \dots, \varepsilon_{i,T})'$,

In this study, the Homogenous Non Causality hypothesis of Dumitrescu and Hurlin (2012) is tested with the null and alternative hypotheses as follows:

$$H_0\colon \gamma_i = 0 \qquad \forall i = 1, \dots, N$$
$$H_1\colon \gamma_i = 0 \qquad \forall i = 1, \dots, N_1;$$
$$\gamma_j \neq 0 \qquad \forall j = N_1 + 1, N_1 + 2, \dots, N$$

Under H_0 it is assumed that there is no causality relationship for all *N*; while under H_1 there are N-N_1 causality relationships, where $N_1 < N$. N_1 is unknown but satisfies the condition $0 \leq N_1/N < 1$. The null hypothesis can be written as $R\theta_i = 0$, where $R = [0\colon I_K]$ is a contrast matrix, constructed by a horizontally concatenated $(K, K+1)$ null matrix 0 and a (K, K) identity matrix I_K, and $\theta_i = (\alpha_i\ \beta_i'\ \gamma_i'\ )'$.

Dumitrescu and Hurlin (2012) show that under H_0 the following panel Wald test statistic $Z_{N,T}^{Hnc}$ will be asymptotically distributed according to a normal distribution with mean zero and variance equals to one as $T \to \infty$:

$$Z_{N,T}^{Hnc} = \sqrt{\frac{N}{2K}}\left(W_{N,T}^{Hnc} - K\right) \qquad (5)$$

where $W_{N,T}^{Hnc} = \frac{1}{N}\sum_{i=1}^{N} W_{i,T}$. $W_{i,T}$ is the individual Wald test statistic for *i*-th cross-section unit corresponding to the individual test $H_0\colon \gamma_i = 0$ and is calculated as follows:

$$W_{i,T} = \hat{\theta}_i' R' [\hat{\sigma}_i^2 R (Z_i' Z_i)^{-1} R']^{-1} R \hat{\theta}_i = \frac{\hat{\theta}_i' R' [R (Z_i' Z_i)^{-1} R']^{-1} R \hat{\theta}_i}{\hat{\varepsilon}_i' \hat{\varepsilon}_i / (T - 2K - 1)} \qquad (6)$$

Where $\hat{\theta}_i$ and $\hat{\varepsilon}_i$ are the OLS estimator for θ_i and the residuals from the regression model (5.4), respectively; $\hat{\sigma}_i^2$ is the variance estimator for the $\hat{\varepsilon}_i$; and $Z_i = [e: Y_i: X_{1i}]$ is a $(T, 2K+1)$ matrix constructed by a horizontally concatenated $(T, 1)$ unit vector, a (T, K) matrix Y_i and a (T, K) matrix X_{1i}.

For a fixed dimension of *T,* normal distribution still holds,[6] however, the panel statistic needs to be standardized and modified to $\tilde{Z}_{N,T}^{Hnc}$ as follows:

$$\tilde{Z}_N^{Hnc} = \sqrt{\frac{N\times(T-2K-5)}{2K\times(T-K-3)}} \times \left[\frac{(T-2K-3)}{(T-2K-1)} \times W_{N,T}^{Hnc} - K\right] \quad (7)$$

In a trivariate setting with an additional explanatory variable X_{2i}, where Y_i, X_{1i}, and X_{2i} are possibly non-stationary stationary variables with different order of integration are described in the following VAR (*K*+*m*) linear model:

$$Y_{i,t} = \alpha_i + \sum_{p=1}^{K+m} \beta_{i,p}\, Y_{i,t-p} + \sum_{p=1}^{K+m} \gamma_{i,p}\, X_{1i,t-p} + \sum_{p=1}^{K+m} \delta_{i,p}\, X_{2i,t-p} + \varepsilon_{i,t} \quad (8)$$

where $Y_i = [y_{i,1}: y_{i,2}: \dots : y_{i,K+m}]'$, $X_{1i} = [x_{1i,1}: x_{1i,2}: \dots : x_{1i,K+m}]'$, and $X_{2i,t} = [x_{2i,1}: x_{2i,2}: \dots : x_{2i,K+m}]'$ are all a $(T, K+m)$ matrix, respectively. All three variables are endogenous with the maximum order of integration *m*. $X_{2i,t}$ is held constant when the Granger causality test $X_{1i,t}$ on $Y_{i,t}$ is conducted. Now, define: the total number of lags $Tlag = K + m$; $Z_i^* = [e: Y_i: X_{1i}: X_{2i}]$ is a $(T, 3Tlag + 1)$ matrix; $\theta_i^* = (\alpha_{1i}\, \beta_{1i}'\, \gamma_{1i}'\, \delta_{1i}')'$ is a $(3Tlag + 1{,}1)$ matrix; $R^* = [0: I_K: 0]$ is a $(K, 3Tlag + 1)$ matrix; $\hat{\theta}_i^*$ and $\hat{\varepsilon}_i^*$ are the OLS estimator for θ_i and the residuals from the regression model (8), respectively; and $\hat{\sigma}_i^{*2}$ is the variance estimator for the $\hat{\varepsilon}_i^*$.

Using the fact that the rank of R^* is still the same with that of *R*, the Dumitrescu and Hurlin (2012) panel non-causality test in heterogeneous panels still can be applied by modifying the Wald statistics of $W_{i,T}$, $Z_{N,T}^{Hnc}$, and $\tilde{Z}_{N,T}^{Hnc}$ with:[7]

[6] Dumitrescu and Hurlin (2012) also formulate approximated critical values for fixed *N* and *T* samples. However, their Monte Carlo simulation provides evidence that the standardized $\tilde{Z}_{N,T}^{Hnc}$ also performs well when *N* is small as in our case.

[7] See Appendix for proofs.

$$W_{i,T}^{*} = \hat{\theta}_i^{*\prime} R^{*\prime}[\hat{\sigma}_i^{*2} R^{*}(Z_i^{*\prime} Z_i^{*})^{-1} R^{*\prime}]^{-1} R^{*} \hat{\theta}_i^{*} = \frac{\hat{\theta}_i^{*\prime} R^{*\prime}[R^{*}(Z_i^{*\prime} Z_i^{*})^{-1} R^{*\prime}]^{-1} R^{*} \hat{\theta}_i^{*}}{\hat{\varepsilon}_i^{*\prime}\hat{\varepsilon}_i^{*}/(T-3Tlag-1)} \quad (9)$$

$$Z_{N,T}^{Hnc*} = \sqrt{\frac{N}{2K}}(W_{N,T}^{Hnc*} - K) \quad (10)$$

$$\tilde{Z}_{N}^{Hnc*} = \sqrt{\frac{N\times(T-3Tlag-5)}{2K\times(T-2K-3m-6)}} \times \left[\frac{(T-3Tlag-3)}{(T-3Tlag-1)} \times W_{N,T}^{Hnc*} - K\right] \quad (11)$$

To accommodate cross-sectional dependence, Dumitrescu and Hurlin (2012) propose using bootstrapped critical values. This study adapts their bootstrapping technique to the trivariate Toda and Yamamoto (1995) framework in the following steps:[8]

1. Estimate model (8) under the null hypothesis, that is set $\gamma_{i,p} = 0, \forall p = 1, \dots, K$ for all i and obtain the residuals;
2. Resample the residuals by choosing a complete row in the residual matrix to preserve the cross-correlation structure;
3. Construct a resampled series $y_{i,t}$ under the null hypothesis i.e. $y_{i,t}^{*} = \hat{\alpha}_i + \sum_{p=1}^{K+m} \hat{\beta}_{i,p}\, y_{i,t-p} + \sum_{p=K+1}^{m} \hat{\gamma}_{i,p}\, X_{1i,t-p} + \sum_{p=1}^{K+m} \hat{\delta}_{i,p}\, X_{2i,t-p} + \hat{\varepsilon}_{i,t}$ and compute the Wald statistics;
4. Repeat steps 2 and 3 many times to construct a series of the Wald statistics. Select the appropriate percentiles of the series to recover bootstrapped critical values.

3. Data

The proxy for portfolio equity flow data is the net portfolio investment of equity (in millions USD) collected from the balance of payment statistics (under BPM5) of the IFS published by the International Monetary Funds. Net portfolio inflows are then calculated by subtracting assets from liabilities of the net portfolio investments and expressed as percentage of current GDP.[9] The MSCI series for the end of period

[8] The Matlab code used here for trivariate Granger non-causality tests builds on the programs provided by Hurlin (http://www.runmycode.org/companion/view/42) and by Emirmahmutoglu (http://www.runmycode.org/companion/view/89). The code is available upon request.

[9] As an alternative, one may modify the international financial integration measure of Lane and Milesi-Ferretti (2007) by only using the assets and liabilities of *equity securities* to measure international *equity* integration.

exchange rates per US dollar and the stock indices are collected from Thomson Reuters Datastream Professional.[10]

This study are interested at exchange rate dynamics, therefore the initial sampling frame is all countries that implement managed or free float exchange rate arrangement. Those adopting Euro as their official currency are excluded from the sample. As Indonesia is the focus in the thesis, it then is included. The study period is therefore chosen by using Indonesia as a benchmark for determining the longest data series. The final dataset comprises of eight economies covering both advanced and emerging markets implementing managed or free float exchange rate arrangement. They are Australia, Canada, Indonesia, Japan, South Korea, Sweden, Thailand and U.K. The sample period of 1993:Q1-2008:Q4 is chosen as the one that ensures the longest available quarterly series.[11] Our sample therefore consists of eight cross section units and 64 time series units ($N = 8$ and $T = 64$). All series are not seasonally adjusted data. They all are expressed in a natural logarithm, except the net portfolio capital inflows that may contain a negative value reflecting capital outflows.

This paper examines both individual time series data and panel data. The unit root tests of Zivot and Andrews (2002) and the cointegration test of Gregory and Hansen (1996) are applied to the individual time series data. These tests allow for the presence of a single structural break in the time series. For the panel data, this study employs the Pesaran (2004) tests for cross section dependence and the modified Sargan-Bhargava (MSB) panel unit root test of Bai and Carrion-i-Silvestre (2009). The panel unit root test is of the so-called third generation of panel unit root tests which use common factors to represent cross-sectional dependence and allow for the presence of unknown multiple structural breaks at different dates. It can also detect the breaks when they exist. The panel cointegration test of Banerjee and Carrion-i-

[10] It is interesting to note that different studies may use different forms of data either in level (prices or rates) or in first difference (rate of returns). For instance, Hau and Rey (2005), Inci and Lee (2014), Yang *et al*. (2014), Caporale *et al*. (2014), use rates of returns; while Granger *et al.* (2000), Tsagkanos and Siriopoulos (2013), Moore and Wang (2014) and Groenewold and Paterson (2013) use prices and exchange rates. The common approach is that if a unit root test is failed to be rejected for first differenced data, then the rate of returns is used. However, as explained in the methodology section, this approach may be misleading when Granger causality test is employed.

[11] The longer sample period is available for the balance of payment statistics under BPM6 of the IFS published by the International Monetary Funds. However, there are changes in treatment and classification between BPM5 and BPM6. Equity securities in BMP5 exclude investment fund shares, while in BPM6 investment fund shares are included, which in the authors' opinion do not fully reflect portfolio equity.

Silvestre (2015) that allows for structural breaks and cross-section dependence is also employed.

4. Results

4.1 Descriptive Statistics

The individual country time-series plots for all variables of interest are presented in Figure 2, Figure 3 and Figure 4. Australia, Sweden, and U.K. are net recipients of capital inflows over the period of study with an average around 2.5-3.5 of GDP. Indonesia was hit hard by the 1997 Asian financial crisis which caused massive capital outflows in the fourth quarter of 1997 to the second quarter of 1998. Figure 2 shows that portfolio equity flows were relatively stable before the crisis and have been relatively more volatile since then. An increasing trend in stock prices and exchange rates is a common feature in all economies; while the variable net capital inflows of portfolio equity fluctuate around zero over the period of 1999-2001. The volatility levels of stock prices and equity flows are relatively similar, but they are consistently higher than that of exchange rates as shown in Table 1.

INSERT FIGURE 2 ABOUT HERE

INSERT FIGURE 3 ABOUT HERE

INSERT FIGURE 4 ABOUT HERE

INSERT TABLE 1 ABOUT HERE

The Pearson product-moment correlation coefficients indicate a negative association between stock prices and exchange rates as well as between exchange rates and portfolio equity flows. The degree of association for the latter is, however, weaker than that of the former. Meanwhile, a positive association exists between stock prices and portfolio equity inflows. Table 1 also indicates the presence of cross-country dependence among economies in the sample, which may be due to the financial market integration or spill over effects between countries. The average cross-sectional dependence correlation coefficients for all variables are positive and statistically significant at one per cent level.

4.2 Individual Time Series

Table 2 presents the result of Zivot and Andrews (2002)'s unit root tests that allows for a single break. In general, the tests show that $EqFlows_{i,t}$ is I(0), while $Index_{i,t}$ and $Currency_{i,t}$ are I(1). However, for Indonesian Rupiah and Thailand Baht, the results indicate that they could be I(0). Based on these unit root tests, there is a need to examine the cointegration relationship between the two I(1) processes. Cointegration tests here are needed for verification, but they do not affect the Toda-Yamamoto test. Table 3 presents the result of Gregory and Hansen (1996)'s cointegration tests with regime shift for the variables. When $Index_{i,t}$ is regressed on $Currency_{i,t}$ – as in our main interest – all countries except Korea and possibly Thailand show that there is no cointegration relationship between the variables. However when $Currency_{i,t}$ is regressed on $Index_{i,t}$, the tests indicate cointegration for Indonesia.

INSERT TABLE 2 ABOUT HERE

INSERT TABLE 3 ABOUT HERE

Table 4 shows that the maximum lag length to be used in a standard VAR model vary, depending on the criteria used. The three criteria, i.e. Akaike's information criterion (AIC), Schwarz's Bayesian information criterion (SBIC), and the Hannan and Quinn information criterion (HQIC), indicate that the maximum lag length generally varies from one to two. Only Indonesia is indicated to have the maximum number of lag of four. Based on these criteria, it is concluded that the maximum lag is either one ($K = 1$) or two ($K = 2$).

INSERT TABLE 4 ABOUT HERE

The results for Granger causality test in a Toda-Yamamoto framework and the signs of the first lag parameter estimate for the independent variable of interest are summarized in Table 5. Column (1) in Panel A shows that stock prices Granger cause exchange rates in cases of Korea, Thailand and UK as indicated by the individual Wald statistics for these economies that are statistically significant at 5 per cent level. The signs of the parameter estimate for the first lag of stock price variable ($Index_{i,t-1}$) are negative as predicted in the portfolio balance approach models. The presence of a cointegration relation between the two variables in Table 3 also confirms these causality test results, at least for Korea and Thailand. However, column (4) shows no evidence for the risk rebalancing channel for portfolio equity flows because there is no such case where portfolio equity flows Grange cause exchange rates in those three countries. For Indonesia, Korea and UK, portfolio equity flows Grange cause exchange rates. However, an evidence for that portfolio equity flows is Granger caused by stock prices only exists for Indonesia (column 2). The sign of the parameter estimate for $Index_{i,t-1}$ is also negative.

Panel B presents the results in case K = 2. The Granger causality between stock prices and exchange rates still exists for Thailand and UK, but not for Indonesia. However, portfolio equity flows still Granger cause exchange rates in case of Indonesia with a negative parameter estimate of $EqFlows_{i,t-1}$. In contrast to panel A, panel B shows that a causality from exchange rates to stock prices may exist for Indonesia, Korea and Japan. In general Table 5 shows that the portfolio balance hypothesis, in particular the risk rebalancing channel for portfolio equity flow approach, is only supported in the case of Indonesia. In this case, stock prices positively affect portfolio equity flows, and conversely, portfolio equity flows negatively affects exchange rates.

INSERT TABLE 5 ABOUT HERE

Compared to the results of bivariate analysis of other studies, this trivariate study provides similar findings. Similar to the findings of Hau and Rey (2005) and Groenewold and Paterson (2013), no evidence of Granger causality was found between stock prices and exchange rates for Australia. Similarly, no evidence of Granger causality was found in Japan, consistent with Granger *et al.* (2000), Hau and Rey (2005), and Caporale *et al.* (2014). A unidirectional causality from stock prices to exchange rates for UK found was found in this study similar to the findings of Hau and Rey (2005) and Caporale *et al.* (2014). In case of Korea and Thailand this study supports the feedback relations as was found by Andreou *et al.* (2013) and Yang *et al.* (2014). However, unlike Caporale *et al.* (2014) and Hau and Rey (2005) that found a causality relation between stock prices and exchange rates, this study fails to find such a relationship for Canada and Sweden. In particular for Indonesia, this study's finding may resolve conflicting findings from other studies. Studies by Andriansyah (2003) and Lee *et al.* (2011) provide evidence for stock prices Granger cause exchange rates, Liang *et al.* (2013) on the other hand support the reverse causality direction. Bi-directional causality for Indonesia is supported by Yang *et al.* (2014), while no evidence for causality is provided by Granger *et al.* (2000).

4.3 Panel Data

The MSB test of Bai and Carrion-i-Silvestre (2009) provides three different panel statistics and their corresponding simplified statistics. In case of no structural breaks, the panel and simplified statistics produce the same values. The first statistic is Z^*, the average of individual statistics which follows the standard normal distribution. The other statistics are P^* and P_m^*, the average of individual p-values. P^*-statistic is designed for a fixed number of cross-sections, while P_m^*-statistic is designed for large number of cross-sections. As our sample has a limited number of cross-sections, we are more interested at P^*-statistic. The simplified statistics as shown in Table 6 indicate that both exchange rates and stock prices contain a unit root, while portfolio equity does not. The panel unit root test also shows no evidence for any structural break in our series. To check robustness of the results of the unit root test, we also employ the cross-sectionally augmented Dickey-Fuller (CADF) test of Pesaran (2007) and the cross-sectionally augmented Sargan-Bhargava (CSB) test of Pesaran *et al.* (2013). Both tests confirm that $Currency_{i,t}$ and $Index_{i,t}$ are I(1) processes, and $EqFlows_{i,t}$ is I(0) process (see Table 7).

INSERT TABLE 6 ABOUT HERE

INSERT TABLE 7 ABOUT HERE

The next step is to examine the possibility of cointegration relationship between I(1) series: $Currency_{i,t}$ and $Index_{i,t}$. As an alternative for Banerjee and Carrion-i-Silvestre (2015) test, this study also employs the panel cointegration tests of Westerlund (2007) and Di Iorio and Fachin (2014). These tests apply the residual-based stationary bootstrap test to account for cross-section dependence. In terms of small sample properties, Di Iorio and Fachin (2014) claim that their test is preferable to the other panel cointegration tests. Table 8 summarizes the three panel cointegration tests which provide insufficient evidence for cointegration. All test statistics cannot reject the null hypothesis of no-cointegration, except for the G_t statistic of Westerlund (2007).

Based on the above results, the number of additional lags is set to one (m = 1) and the order of panel VAR is set according to the results from the individual time series, i.e. either $K = 1$ or $K = 2$.[12] The results of the trivariate Toda-Yamamoto approach for Granger non-causality test in heterogeneous panels are summarized in Table 9 below.

INSERT TABLE 8 ABOUT HERE

INSERT TABLE 9 ABOUT HERE

Similar to individual time-series, Table 9 provides no evidence for the risk rebalancing channel for portfolio equity flow approach in the panel data setting. In general stock prices Granger cause exchange rates and portfolio equity flows Granger cause exchange rates. However, there is no evidence that stock prices Granger cause portfolio equity flows which is necessary to support the portfolio balance approach as illustrated in Figure 1.

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[12] A Stata command called *pvarsoc* provides lag-order selection statistics for panel VAR estimated using GMM. It reports MMSC-Bayesian information criterion, MMSC-Akaike's information criterion, and MMSC-Hannan and Quinn information criterion. Using this command, the recommended the value for K is 1.

4.4 Panel Data: Pre- and Post-Crisis Periods

It is well known that financial crises affect both the financial and economic sectors, such as asset prices, output, and employment (Reinhart & Rogoff 2009). Moreover, it is plausible that unusual shocks may give rise to structural shifts or asymmetries. For example, Evgenidis and Tsagkanos (2017) find that negative shocks in the post-crisis Great Recession period have impacted on the real economy at a greater extent than positive shocks in the transmission mechanism. Li (2013) also finds asymmetric co-movements between the U.S. stock market and some developed stock markets where market downturns lead to stronger co-movements than market upturns.

Although asymmetries or structural changes are potentially important, the linear VAR models employed in this paper cannot accommodate asymmetries or structural breaks caused the period of crisis. A threshold-VAR/ECM approach employed by Evgenidis and Tsagkanos (2017) and Evgenidis et al. (2017) could be an alternative. They find asymmetries using time series data up to 2013. However, due to lack of availability, our *panel* data only covers the period up to 2008 and thus we do not have enough post-crisis data to account for possible asymmetries in the 2008 financial crisis period. Thus, future work may explore longer time-series and structural breaks or asymmetries as well as the possibility of a *panel* threshold VAR model.

To isolate the impact of the Asian Financial crisis happened over the period 1997:Q2-1998:Q4, this paper rather re-estimates the trivariate Toda-Yamamoto approach for Granger non-causality test in heterogeneous panels by splitting the sample period in two sub-periods: pre- and post-crises (i.e., 1997 Asian financial crisis). The result for the pre-crisis period (1993:Q1-1997Q1) is presented in Table 10, while that for the post-crisis period (1999:Q1-2008:Q4) is in Table 11.

The finding for both sub-periods are similar to the general finding which is there is no evidence to support the portfolio balance approach. For the pre-crisis period, stock prices Granger still cause exchange rates. Portfolio equity flows also still Granger cause exchange rates, even though these results are not as strong as before. For the post-crisis, the result the portfolio balance approach is even not supported by the fact that stock prices do not statistically Granger cause exchange rates. The reverse, however, is still the case. Exchange rates Granger cause stock prices.

INSERT TABLE 10 ABOUT HERE

INSERT TABLE 11 ABOUT HERE

5. Conclusion

This study re-examines a portfolio model prediction of a negative causal relationship between stock prices and exchange rates through portfolio capital flow transmission channel. The bivariate stationary Granger non-causality test in heterogeneous panels of Dumitrescu and Hurlin (2012) is extended to a trivariate setting in the framework of Toda and Yamamoto (1995). The variables of interest in this framework may be non-stationary and integrated at different order. This study uses a macro panel data setting for eight emerging and developed economies with managed or fee floating exchange rate arrangement. The evidence suggest that stock prices Granger cause exchange rates and portfolio equity flows Granger cause exchange rates. However, the overall panel evidence casts doubt on the explicit trivariate model of portfolio balance model examined here. In our panel study, only in Indonesia stock prices affect exchange rates via the portfolio equity flow channel.

It is, however, important to note again that our paper has utilised panel data that has limited our exploration into non-linearities in the above trivariate relationship. It is theoretically possible that the results reported in this paper may have been influenced by non-linearities that could not be considered here. It is well known that the presence of asymmetries or structural breaks can cause estimation problems, such as biased coefficient estimates, if they are not accounted for. Hence, future research ought to more comprehensively examine the potential effects of asymmetries or structural shifts in the trivariate relationship of interest here.

The evidence from this study has important implications for policy makers and investors in understanding the relationship between the three variables considered here. First, it is important to be informed of the factors that impact on the exchange rate in order to facilitate exchange rate stability. Second, it is also imperative to identify causal effects between the monetary sector and the real economy, in particular on investment and GDP growth. For instance, governments may wish to integrate capital market with fiscal policies to maintain a stable exchange rate. Finally, investors may benefit from insights on the causal links relating to exchange rate movements when they make decisions on international portfolio management.

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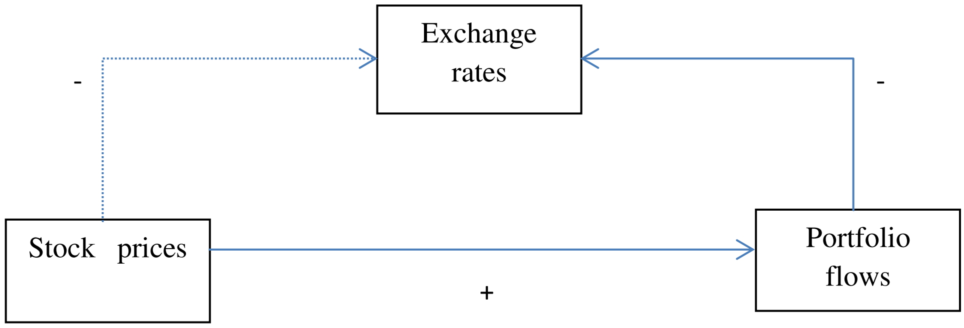

Figure 1. Expected Relationship between Stock Prices, Exchange Rates and Portfolio Equity Flows

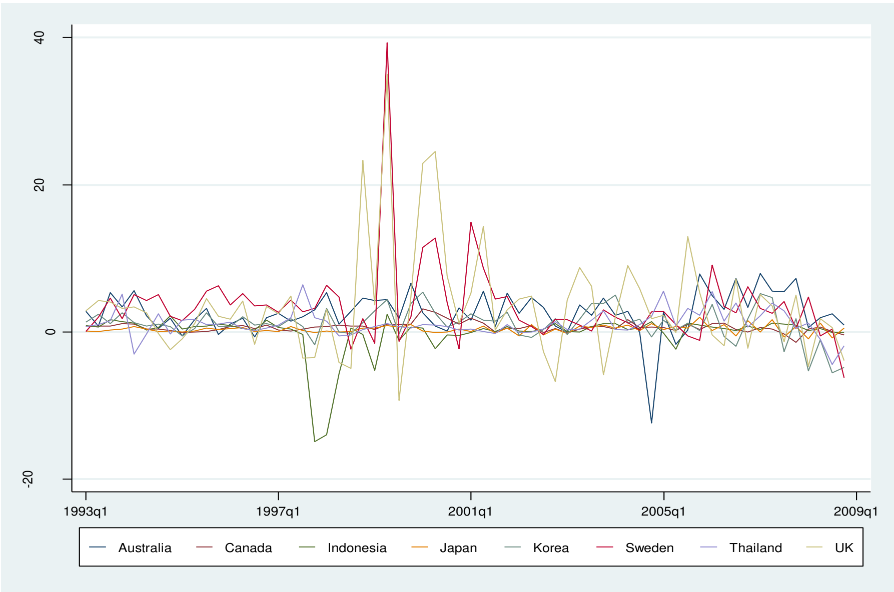

Figure 2. Portfolio Equity Flows (as percentage of GDP)

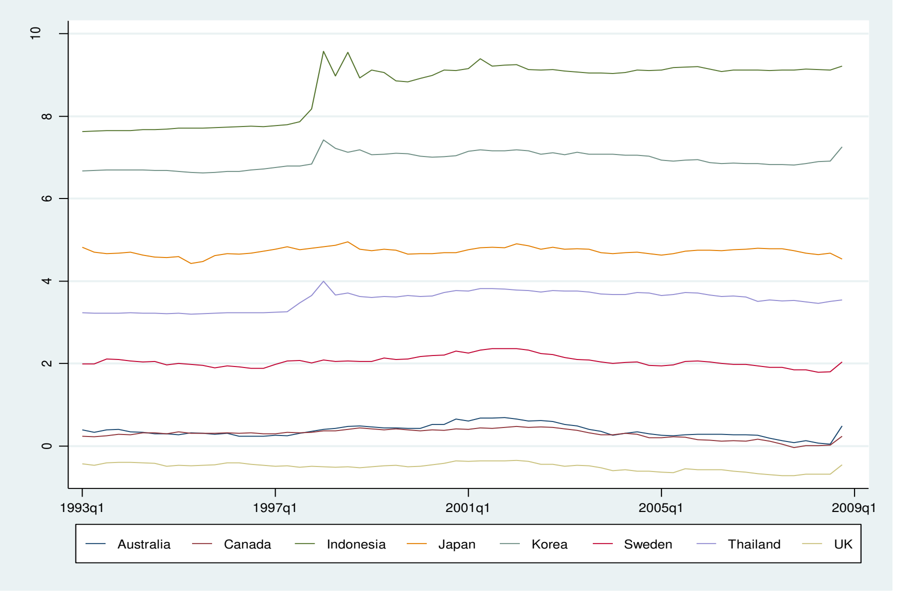

Figure 3. Exchange Rates (in natural logarithm)

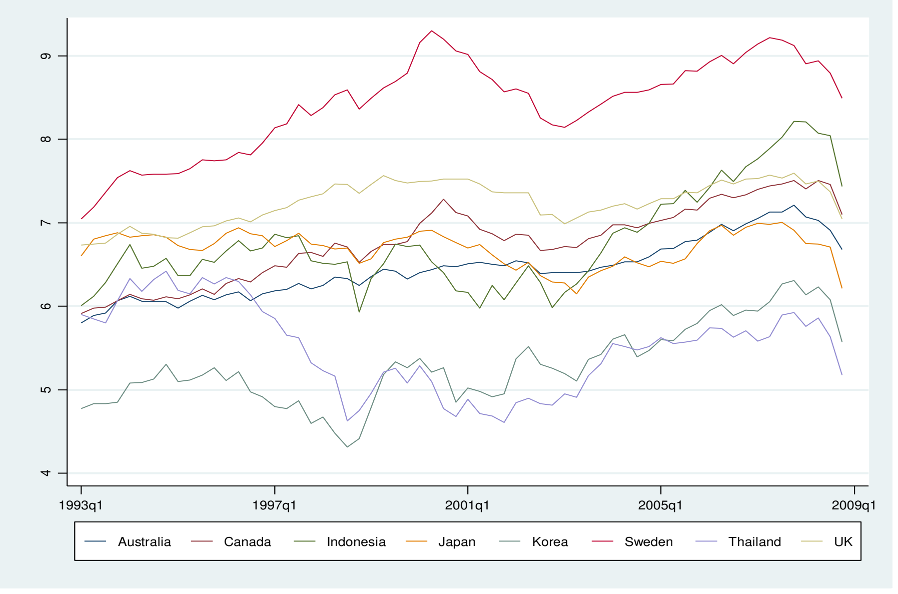

Figure 4. Stock Prices (in natural logarithm)

Table 1. Descriptive Statistics (for Panel Data)

Statistics	$EqFlows_{i,t}$	$Currency_{i,t}$	$Index_{i,t}$
No. observations	512	512	512
Mean	1.578	3.262	6.645
Median	0.864	2.784	6.670
St.dev	3.989	3.138	1.018
Min	39.283	9.571	9.299
Max	-14.899	-0.719	4.312
Pearson correlation			
Capital flows	1.000		
Currency	-0.229***	1.000	
Index	0.192***	-0.341***	1.000
Pesaran (2004) test for cross-sectional independence			
Averaged correlation coefficient	0.162	0.460	0.405
CD-statistic	6.84***	19.49***	17.13***

Notes: The null hypothesis of cross-section independence CD-statistic follows a standard normal distribution. All correlation coefficients and CD-statistics are significant at 1 per cent level (denoted by ***).

Table 2. The Unit Root Tests of Zivot and Andrews (2002)

	$EqFlows_{i,t}$	$Currency_{i,t}$	$Index_{i,t}$	$\Delta Currency_{i,t}$	$\Delta Index_{i,t}$
Australia	-7.464***	-3.201	-2.594	-5.333**	-6.707***
Canada	-5.113**	-3.396	-3.202	-5.278**	-7.041***
Indonesia	-4.950**	-11.015***	-3.369	-5.340***	-8.006***
Japan	-5.080**	-3.788	-2.784	-8.090***	-6.023***
Korea	-6.234***	-4.459*	-3.640	-8.796***	-7.145***
Sweden	-9.777***	-2.929	-2.541	-6.931***	-6.106***
Thailand	-6.143***	-5.622***	-3.259	-9.315***	-7.060***
UK	-9.904***	-2.873	-1.820	-5.757***	-6.920***

Notes: The null hypothesis assumes that all series are non-stationary. The statistics are computed for the model allowing having a break in the intercept. Results are similar when the model allows to have breaks in both the intercept and the slope. The Schwarz Bayesian information criterion is used to decide the number of additional lags. ***, ** and * denote significance at 1 per cent, 5 per cent and 10 per cent level, respectively. The corresponding critical values are -5.34, -4.80, and -4.58, respectively.

Table 3. The Cointegration Tests of Gregory and Hansen (1996)

	ADF	Z_t	Z_a
$Index_{i,t}$ on $Currency_{i,t}$			
Australia	-4.33	-4.12	-25.65
Canada	-3.67	-3.70	-19.84
Indonesia	-4.01	-4.15	-25.35
Japan	-2.25	-2.85	-13.42
Korea	-5.46***	-5.18**	-40.69
Sweden	-3.84	-3.65	-18.49
Thailand	-5.25**	-5.14**	-38.72
UK	-2.97	-2.99	-16.66
$Currency_{i,t}$ on $Index_{i,t}$			
Australia	-3.20	-3.28	-18.89
Canada	-3.39	-3.74	-23.38
Indonesia	-6.86***	-12.66***	-93.56***
Japan	-3.34	-3.83	-22.88
Korea	-7.08***	-7.13***	-62.04***
Sweden	-2.64	-2.82	-13.92
Thailand	-5.72***	-5.15**	-38.58
UK	-3.17	-3.32	-21.15

Notes: The null hypothesis assumes that there is no cointegration between $Index_{i,t}$ and $Currency_{i,t}$. The statistics are computed for the model allowing having a break in the intercept. The Schwarz Bayesian information criterion is used to decide the number of additional lags. ***, ** and * denote significance at 1 per cent, 5 per cent and 10 per cent level, respectively. The corresponding asymptotic critical values are -5.44, -4.92, -4.69; -5.44, -4.92, -4.69; and -57.01, -46.98, -42.49 for *ADF*; Z_t; Z_a; respectively.

Table 4. VAR Lag Order Selection Criteria

	AIC	*HQIC*	*BIC*
Australia	1	1	1
Canada	1	1	1
Indonesia	4	2	2
Japan	2	2	1
Korea	1	1	1
Sweden	1	1	1
Thailand	2	1	1
UK	1	1	1

Notes: The selection of lag order is based on Akaike's information criterion (AIC), Schwarz's Bayesian information criterion (SBIC), and the Hannan and Quinn information criterion (HQIC). The maximum lag is set to four.

Table 5. Trivariate Granger Causality Tests using Toda Yamamoto Framework

	$Index_{i,t}$ → $Currency_{i,t}$	$Index_{i,t}$ → $EqFlows_{i,t}$	$EqFlows_{i,t}$ → $Index_{i,t}$	$EqFlows_{i,t}$ → $Currency_{i,t}$	$Currency_{i}$ → $Index_{i,t}$	$Currency_{i,t}$ → $EqFlows_{i,t}$
	(1)	(2)	(3)	(4)	(5)	(6)
Panel A. K = 1, m = 1						
Australia	1.690	2.069	0.005	0.110	1.448	0.623
	−	+	−	−	+	−
Canada	0.005	1.969	1.471	0.006	0.166	0.676
	−	+	+	+	−	+
Indonesia	1.491	2.975*	0.208	104.022***	1.491	0.047
	−	+	−	−	−	+
Japan	0.989	0.789	9.191***	0.812	0.189	1.559
	−	−	+	+	+	+
Korea	3.300*	1.817	3.741*	5.173**	0.802	0.191
	−	−	+	−	−	−
Sweden	0.435	3.377*	0.003	0.010	0.681	0.468
	−	+	+	+	+	+
Thailand	12.738***	2.718*	8.884***	2.189	0.001	0.769
	−	−	+	+	+	−
UK	5.346**	0.787	0.611	3.112*	0.593	0.308
	−	+	+	+	−	+
Panel B. K = 2, m = 1						
Australia	3.328	1.614	0.430	0.089	1.594	1.002
	−	+	+	+	+	−
Canada	0.691	2.785	2.681	0.202	4.024	1.934
	+	+	+	−	−	+
Indonesia	2.231	2.087	0.132	92.133***	7.069**	0.700
	−	+	−	−	−	−
Japan	0.565	1.531	12.556***	0.569	4.300	2.302
	−	−	+	+	+	+
Korea	4.310	2.243	3.255	4.521	4.875*	0.631
	−	−	+	−	−	−
Sweden	1.144	6.702**	0.154	0.551	0.553	1.323
	−	+	+	+	+	+
Thailand	19.986***	3.149	8.055**	19.199***	8.113**	0.432
	−	−	+	+	−	−
UK	4.948*	1.528	2.286	2.364	0.808	0.379
	−	+	+	+	−	+

Notes: → means the first variable Granger causes the second variable while holding the third variable constant. The null hypothesis assumes that there is no Granger causality from the first variable to the second variable. The individual Wald statistic has a chi-squared distribution with K degrees of freedom. A sign under the Wald statistics indicates the parameter estimate for the first lag of the first variable. ***, ** and * denote significance at 1 per cent, 5 per cent and 10 per cent level, respectively.

Table 6. The MSB Test of Bai and Carrion-i-Silvestre (2009)

Variable	Simplified Test Statistic		
	Z^*	P_m^*	P^*
In levels			
$EqFlows_{i,t}$	-2.921***	29.989***	185.642***
$Currency_{i,t}$	2.182**	-2.112**	4.052
$Index_{i,t}$	-0.334	-0.834	11.280
In first difference			
$\Delta EqFlows_{i,t}$	-2.985***	38.633***	234.542***
$\Delta Currency_{i,t}$	-2.842**	21.678***	138.636***
$\Delta Index_{i,t}$	-2.737***	148.978***	100.731***

Notes: The null hypothesis assumes that all series are non-stationary. The statistics are computed for the model with changes in the slope and allows for maximum two structural changes and maximum six factors. *** and ** denote significance at 1 per cent land 5 per cent level, respectively.

Table 7. The CSB Test of Pesaran *et al*. (2013) and the CADF Test of Pesaran (2007)

Variable	CSB($\hat{p}$) statistic		CADF statistic [Z-t-bar]	
	Lag(1)	Lag(2)	Lag(1)	Lag(2)
In levels				
$EqFlows_{i,t}$	0.036***	0.048***	-6.833***	-4.505***
$Currency_{i,t}$	0.159	0.134	-0.493	-0.208
$Index_{i,t}$	0.198	0.156	2.528	1.872
In first difference				
$\Delta EqFlows_{i,t}$	0.011***	0.013***	-13.207***	12.539***
$\Delta Currency_{i,t}$	0.087***	0.101***	-8.274***	-3.504***
$\Delta Index_{i,t}$	0.026***	0.033***	-7.614***	-5.491***

Notes: The null hypothesis assumes that all series are non-stationary. The statistics are computed by including a linear trend and maximum two lags order. *** denotes significance at 1 per cent level.

Table 8. The Panel Cointegration Tests of Westerlund (2007), Banerjee and Carrion-i-Silvestre (2015) and Di Iorio and Fachin (2014)

Test	Statistic	Critical value/p-value[1]
Banerjee and Carrion-i-Silvestre (2015)[2]		
Z_j^e	0.571	-2.389
		-1.670
		-1.273
Westerlund (2007)[3]		
G_t	-2.855	0.086
G_a	-14.295	0.104
P	-5.873	0.536
P_a	-9.017	0.488
Di Iorio and Fachin (2014)[4]		
Median ADF	-1.989	0.666
Mean ADF	-2.111	0.566
Max ADF	-1.718	0.149

Notes:

[1] The critical values are for Z_j^e statistic at 1 per cent, 5 per cent, and 10 per cent level of significance, respectively. This values are for T=50, the closest number to our sample size. p-values are for the other statistics.

[2] Z_j^e is computed for the individual and time effects model, maximum three number of factors allowed and no structural break. At 5 per cent of significance, 25 per cent of individual tests reject the null hypothesis of no cointegration.

[3] G-statistics are for group mean tests assuming heterogeneity while p-statistics are for the panel test assuming homogeneity. These statistics are computed for the model with constant and trend, maximum two numbers of lags, and the Bartlett kernel window width set of 4. The p-values are robust to cross sectional dependence and computed with 500 bootstrap replications.

[4] ADF statistics are computed for the model with constant and trend, maximum two lags.

Table 9. Trivariate Toda-Yamamoto approach for Granger non-causality test in heterogeneous panels

		Asymptotic Wald Statistics	Bootstrap critical values		
			1%	5%	10%
Panel A. *K = 1, m = 1*					
$Index_{i,t} \rightarrow Currency_{i,t}$	$Z_{N,T}^{Hnc}$	4.308**	5.266	3.531	2.851
	$\tilde{Z}_{N}^{Hnc}$	4.087**	5.011	3.337	2.681
$Index_{i,t} \rightarrow EqFlows_{i,t}$	$Z_{N,T}^{Hnc}$	2.125	5.843	4.043	3.319
	$\tilde{Z}_{N}^{Hnc}$	1.980	5.568	3.831	3.133
$EqFlows_{i,t} \rightarrow Index_{i,t}$	$Z_{N,T}^{Hnc}$	4.028***	1.817	1.474	1.306
	$\tilde{Z}_{N}^{Hnc}$	3.817***	1.683	1.352	1.190
$EqFlows_{i,t} \rightarrow Currency_{i,t}$	$Z_{N,T}^{Hnc}$	26.859***	5.382	4.268	3.651
	$\tilde{Z}_{N}^{Hnc}$	25.846***	5.123	4.048	3.453
$Currency_{i,t} \rightarrow Index_{i,t}$	$Z_{N,T}^{Hnc}$	-0.657	-1.168	-1.099	-1.062
	$\tilde{Z}_{N}^{Hnc}$	-0.704	-1.197	-1.131	-1.095
$Currency_{i,t} \rightarrow EqFlows_{i,t}$	$Z_{N,T}^{Hnc}$	-0.840**	-1.145	-0.722	-0.439
	$\tilde{Z}_{N}^{Hnc}$	-0.881**	-1.175	-0.767	-0.494
Panel B. *K = 2, m = 1*					
$Index_{i,t} \rightarrow Currency_{i,t}$	$Z_{N,T}^{Hnc}$	7.497***	5.862	4.079	3.147
	$\tilde{Z}_{N}^{Hnc}$	3.470***	2.691	1.841	1.396
$Index_{i,t} \rightarrow EqFlows_{i,t}$	$Z_{N,T}^{Hnc}$	2.009	13.359	10.663	9.198
	$\tilde{Z}_{N}^{Hnc}$	0.854	6.265	4.980	4.281
$EqFlows_{i,t} \rightarrow Index_{i,t}$	$Z_{N,T}^{Hnc}$	4.790	8.744	8.089	7.775
	$\tilde{Z}_{N}^{Hnc}$	2.180	4.065	3.753	3.603
$EqFlows_{i,t} \rightarrow Currency_{i,t}$	$Z_{N,T}^{Hnc}$	36.638***	10.906	8.869	7.854
	$\tilde{Z}_{N}^{Hnc}$	17.363***	5.096	4.125	3.641
$Currency_{i,t} \rightarrow Index_{i,t}$	$Z_{N,T}^{Hnc}$	5.422***	1.1578	0.777	0.578
	$\tilde{Z}_{N}^{Hnc}$	2.481***	0.448	0.267	0.1720
$Currency_{i,t} \rightarrow EqFlows_{i,t}$	$Z_{N,T}^{Hnc}$	-2.581**	-2.824	-1.851	-1.255
	$\tilde{Z}_{N}^{Hnc}$	-1.334**	-1.450	-0.986	-0.702

Notes: → means the first variable Granger causes the second variable while holding the third variable constant. The null hypothesis assumes that there is no Granger causality from the first variable to the second variable. The number of iteration for computing bootstrapped critical values is 10,000 times. *** and ** denotes significance at 1 per cent level, and 5 per cent level, respectively.

Table 10. Trivariate Toda-Yamamoto approach for Granger non-causality test in heterogeneous panels (The pre-crisis period)

		Asymptotic Wald Statistics	Bootstrap critical values		
			1%	5%	10%
Panel A. *K = 1, m = 1*					
$Index_{i,t} \rightarrow Currency_{i,t}$	$Z_{N,T}^{Hnc}$	1.939	4.311	3.447	3.175
	$\tilde{Z}_{N}^{Hnc}$	1.151	3.049	2.357	2.140
$Index_{i,t} \rightarrow EqFlows_{i,t}$	$Z_{N,T}^{Hnc}$	1.175*	2.210	1.340	1.036
	$\tilde{Z}_{N}^{Hnc}$	0.540*	1.368	0.672	0.429
$EqFlows_{i,t} \rightarrow Index_{i,t}$	$Z_{N,T}^{Hnc}$	0.692	0.391	0.346	0.322
	$\tilde{Z}_{N}^{Hnc}$	0.154	-0.087	-0.124	-0.143
$EqFlows_{i,t} \rightarrow Currency_{i,t}$	$Z_{N,T}^{Hnc}$	1.067*	1.751	1.266	0.999
	$\tilde{Z}_{N}^{Hnc}$	0.454*	1.001	0.613	0.397
$Currency_{i,t} \rightarrow Index_{i,t}$	$Z_{N,T}^{Hnc}$	-1.020***	1.641	1.755	1.814
	$\tilde{Z}_{N}^{Hnc}$	-1.216***	0.913	1.004	1.052
$Currency_{i,t} \rightarrow EqFlows_{i,t}$	$Z_{N,T}^{Hnc}$	-0.681***	3.131	3.449	3.607
	$\tilde{Z}_{N}^{Hnc}$	-0.945***	2.105	2.359	2.486
Panel B. *K = 2, m = 1*					
$Index_{i,t} \rightarrow Currency_{i,t}$	$Z_{N,T}^{Hnc}$	17.561***	16.378	15.107	14.417
	$\tilde{Z}_{N}^{Hnc}$	4.732***	4.366	3.973	3.759
$Index_{i,t} \rightarrow EqFlows_{i,t}$	$Z_{N,T}^{Hnc}$	9.474	12.348	11.342	10.935
	$\tilde{Z}_{N}^{Hnc}$	2.230	3.119	2.808	2.682
$EqFlows_{i,t} \rightarrow Index_{i,t}$	$Z_{N,T}^{Hnc}$	33.459***	10.578	10.263	10.125
	$\tilde{Z}_{N}^{Hnc}$	9.649***	2.572	2.474	2.432
$EqFlows_{i,t} \rightarrow Currency_{i,t}$	$Z_{N,T}^{Hnc}$	0.496	3.080	2.574	2.302
	$\tilde{Z}_{N}^{Hnc}$	-0.546**	-0.580	-0.494	-0.440
$Currency_{i,t} \rightarrow Index_{i,t}$	$Z_{N,T}^{Hnc}$	19.357	20.797	20.416	20.181
	$\tilde{Z}_{N}^{Hnc}$	5.287	5.733	5.615	5.542
$Currency_{i,t} \rightarrow EqFlows_{i,t}$	$Z_{N,T}^{Hnc}$	11.228	35.222	19.540	14.923
	$\tilde{Z}_{N}^{Hnc}$	2.773	10.194	5.344	3.916

Notes: → means the first variable Granger causes the second variable while holding the third variable constant. The null hypothesis assumes that there is no Granger causality from the first variable to the second variable. The number of iteration for computing bootstrapped critical values is 10,000 times. *** and ** denotes significance at 1 per cent level, and 5 per cent level, respectively.

Table 11. Trivariate Toda-Yamamoto approach for Granger non-causality test in heterogeneous panels (The post-crisis period)

		Asymptotic Wald Statistics	Bootstrap critical values		
			1%	5%	10%
Panel A. *K = 1, m = 1*					
$Index_{i,t} \rightarrow Currency_{i,t}$	$Z_{N,T}^{Hnc}$	2.637	5.945	4.196	3.542
	$\tilde{Z}_{N}^{Hnc}$	2.356	5.464	3.821	3.207
$Index_{i,t} \rightarrow EqFlows_{i,t}$	$Z_{N,T}^{Hnc}$	1.490	3.395	2.281	1.773
	$\tilde{Z}_{N}^{Hnc}$	1.279	3.068	2.022	1.544
$EqFlows_{i,t} \rightarrow Index_{i,t}$	$Z_{N,T}^{Hnc}$	5.185*	5.685	5.261	5.003
	$\tilde{Z}_{N}^{Hnc}$	4.750*	5.220	4.821	4.578
$EqFlows_{i,t} \rightarrow Currency_{i,t}$	$Z_{N,T}^{Hnc}$	1.527	11.326	10.176	9.704
	$\tilde{Z}_{N}^{Hnc}$	1.314	10.519	9.438	8.995
$Currency_{i,t} \rightarrow Index_{i,t}$	$Z_{N,T}^{Hnc}$	-0.244***	0.006	0.125	0.193
	$\tilde{Z}_{N}^{Hnc}$	-0.350***	-0.116	-0.004	0.060
$Currency_{i,t} \rightarrow EqFlows_{i,t}$	$Z_{N,T}^{Hnc}$	-0.133**	-0.383	0.021	0.284
	$\tilde{Z}_{N}^{Hnc}$	-0.246**	-0.481	-0.101	0.146
Panel B. *K = 2, m = 1*					
$Index_{i,t} \rightarrow Currency_{i,t}$	$Z_{N,T}^{Hnc}$	3.5175	8.421	6.766	5.918
	$\tilde{Z}_{N}^{Hnc}$	1.426	3.671	2.917	2.525
$Index_{i,t} \rightarrow EqFlows_{i,t}$	$Z_{N,T}^{Hnc}$	1.333	16.495	13.426	12.133
	$\tilde{Z}_{N}^{Hnc}$	0.425	7.369	5.964	5.371
$EqFlows_{i,t} \rightarrow Index_{i,t}$	$Z_{N,T}^{Hnc}$	4.751	12.150	11.564	11.299
	$\tilde{Z}_{N}^{Hnc}$	1.991	5.379	5.110	4.989
$EqFlows_{i,t} \rightarrow Currency_{i,t}$	$Z_{N,T}^{Hnc}$	0.150	22.226	19.917	18.898
	$\tilde{Z}_{N}^{Hnc}$	-0.117***	4.964	5.551	5.876
$Currency_{i,t} \rightarrow Index_{i,t}$	$Z_{N,T}^{Hnc}$	0.2890	6.087	5.754	5.574
	$\tilde{Z}_{N}^{Hnc}$	-0.052***	1.595	1.760	1.8230
$Currency_{i,t} \rightarrow EqFlows_{i,t}$	$Z_{N,T}^{Hnc}$	0.480	14.419	11.794	10.740
	$\tilde{Z}_{N}^{Hnc}$	0.035	6.418	5.216	4.733

Notes: → means the first variable Granger causes the second variable while holding the third variable constant. The null hypothesis assumes that there is no Granger causality from the first variable to the second variable. The number of iteration for computing bootstrapped critical values is 10,000 times. *** and ** denotes significance at 1 per cent level, and 5 per cent level, respectively.

APPENDIX

Proofs of Equations (9), (10), and (11)

Equation (9)

In the context of bivariate setting, by defining $\widetilde{\varepsilon_i} = \varepsilon_i/\sigma_{\varepsilon i}$, an individual Wald test $W_{i,T}$ can be expressed in the form of

$$W_{i,T} = (T - 2K - 1)\left(\frac{\widetilde{\varepsilon_i}'\Phi_i\widetilde{\varepsilon_i}}{\widetilde{\varepsilon_i}'M_i\widetilde{\varepsilon_i}}\right)$$

Dumitrescu and Hurlin (2012) argue that $W_{i,T}$ has the same chi-square distribution as $\widetilde{\varepsilon_i}'\Phi_i\widetilde{\varepsilon_i}$ with a degree of freedom equal to the rank of Φ_i. They further show that the rank of $\Phi_i$is the same as the rank of *R* which is *K*.

After adjusting the definitions of matrices $\hat{\theta}_i, R, Z_i$ and $\hat{\varepsilon}_i$ in bivariate setting into their trivariate setting $\hat{\theta}_i^*, R^*, Z_i^*$ and $\hat{\varepsilon}_i^*$, a modified Dumitrescu and Hurlin (2012)'s individual Wald test $W_{i,T}^*$ can be calculated as follows:

$$W_{i,T}^* = \hat{\theta}_i^{*\prime}R^{*\prime}[\hat{\sigma}_i^{*2}R^*(Z_i^{*\prime}Z_i^*)^{-1}R^{*\prime}]^{-1}R^*\hat{\theta}_i^* = \frac{\hat{\theta}_i^{*\prime}R^{*\prime}[R^*(Z_i^{*\prime}Z_i^*)^{-1}R^{*\prime}]^{-1}R^*\hat{\theta}_i^*}{\hat{\varepsilon}_i^{*\prime}\hat{\varepsilon}_i^*/(T-3Tlag-1)}$$

Following the same logic above, $W_{i,T}^*$ can similarly be expressed in the form of

$$W_{i,T}^* = (T - 3(K+m) - 1)\left(\frac{\widetilde{\varepsilon_i^{*}}'\Phi_i\widetilde{\varepsilon_i^*}}{\widetilde{\varepsilon_i^{*}}'M_i\widetilde{\varepsilon_i^*}}\right)$$

and $W_{i,T}^*$ will have the same chi-square distribution as $\widetilde{\varepsilon_i^{*}}'\Phi_i\widetilde{\varepsilon_i^*}$ with a degree of freedom equal to the rank of R^*. Because $R^* = [0{:}\, I_K{:}\, 0]$, its rank will be the same as $R = [0{:}\, I_K]$: that is *K*. Therefore, when $T \to \infty$, $W_{i,T}^* \xrightarrow{d} \chi^2(K), \forall_i = 1, \ldots, N$ still holds.

Equation (10)

In addition, when $N \to \infty$, $E\left(W_{i,T}^{*}\right) = K$ and $Var\left(W_{i,T}^{*}\right) = 2K$, the Linderberg-Levy central limit theorem conjectures that $\sqrt{N}\left(\frac{1}{N}\sum_{i=1}^{N} W_{i,T}^{*} - K\right) \xrightarrow{d} N(0,2K)$. After a normalization, it can be shown that $Z_{N,T}^{Hnc*} = \sqrt{\frac{N}{2K}}(W_{N,T}^{Hnc*} - K)$, then $Z_{N,T}^{Hnc*} \to N(0,1)$.

Equation (11)

Dumitrescu and Hurlin (2012) show that the statistic needs to be adjusted for a fixed T sample. Because the rank of R^{*} is still K, the moment of individual Wald can be modified as follows:

$$N^{-1}\sum_{t=1}^{N} E(W_{i,T}^{*}) \cong E\left(\widetilde{W}_{i,T}^{*}\right) = K \times \frac{(T - 3Tlag - 1)}{(T - 3Tlag - 3)}$$

With the second moment, $E\left[\left(\widetilde{W}_{i,T}^{*}\right)^{2}\right] = \frac{(T-3Tlag-1)^{2}\times(2K+K^{2})}{(T-3Tlag-3)\times(T-3Tlag-5)}$, its variance can be calculated as follows: $Var\left(\widetilde{W}_{i,T}^{*}\right) = 2K \times \frac{(T-3Tlag-1)^{2}\times(T-2K-3m-6)}{(T-3Tlag-3)^{2}\times(T-3Tlag-5)}$

Proof:

$$Var\left(\widetilde{W}_{i,T}^{*}\right) = E\left[\left(\widetilde{W}_{i,T}^{*}\right)^{2}\right] - \left[E\left(\widetilde{W}_{i,T}^{*}\right)\right]^{2}$$

$$= \frac{(T - 3Tlag - 1)^{2} \times (2K + K^{2})}{(T - 3Tlag - 3) \times (T - 3Tlag - 5)} - \left[K \times \frac{(T - 3Tlag - 1)}{(T - 3Tlag - 3)}\right]^{2}$$

$$= \frac{[(T - 3Tlag - 1)^{2} \times (2K + K^{2}) \times (T - 3Tlag - 3)] - [K^{2} \times (T - 3Tlag - 1)^{2} \times (T - 3Tlag - 5)]}{(T - 3Tlag - 3)^{2} \times (T - 3Tlag - 5)}$$

The denominator can be simplified as follows:

$$= (T - 3Tlag - 1)^{2} \times [(2K + K^{2}) \times (T - 3Tlag - 3) - K^{2} \times (T - 3Tlag - 5)]$$

$$= (T - 3Tlag - 1)^{2} \times [2KT - 6KTlag - 6K + K^{2}T - 3K^{2}Tlag - 3K^{2} - K^{2}T + 3K^{2}Tlag + 5K^{2}]$$

$$= (T - 3Tlag - 1)^{2} \times [2KT - 6KTlag - 6K + 2K^{2}]$$

$$= 2K \times (T - 3Tlag - 1)^2 \times [T - 3Tlag - 6 + K]$$
$$= 2K \times (T - 3Tlag - 1)^2 \times [T - 3K - 3m - 6 + K]$$
$$= 2K \times (T - 3Tlag - 1)^2 \times [T - 2K - 3m - 6]$$

Therefore, $Var(\widetilde{W}_{i,T}^{*}) = 2K \times \frac{(T-3Tlag-1)^2 \times (T-2K-3m-6)}{(T-3Tlag-3)^2 \times (T-3Tlag-5)}$

Meanwhile,

$$\tilde{Z}_N^{Hnc*} = \frac{\sqrt{N}\left[W_{N,T}^{Hnc*} - E(\widetilde{W}_{i,T}^{*})\right]}{\sqrt{Var(\widetilde{W}_{i,T}^{*})}} = \frac{\sqrt{N}\left[W_{N,T}^{Hnc*} - K \times \frac{(T-3Tlag-1)}{(T-3Tlag-3)}\right]}{\sqrt{2K \times \frac{(T-3Tlag-1)^2 \times (T-2K-3m-6)}{(T-3Tlag-3)^2 \times (T-3Tlag-5)}}}$$

$$= \frac{\sqrt{N}\left[W_{N,T}^{Hnc*} - K \times \frac{(T-3Tlag-1)}{(T-3Tlag-3)}\right]}{\frac{(T-3Tlag-1)}{(T-3Tlag-3)}\sqrt{2K \times \frac{(T-2K-3m-6)}{(T-3Tlag-5)}}}$$

$$= \frac{\sqrt{N}\left[W_{N,T}^{Hnc*} - K \times \frac{(T-3Tlag-1)}{(T-3Tlag-3)}\right] \times (T-3Tlag-3)}{(T-3Tlag-1) \times \sqrt{2K \times \frac{(T-2K-3m-6)}{(T-3Tlag-5)}}}$$

$$= \frac{(T-3Tlag-3)}{(T-3Tlag-1)} \times \sqrt{\frac{N \times (T-3Tlag-5)}{2K \times (T-2K-3m-6)}}\left[W_{N,T}^{Hnc*} - K \times (T-3Tlag-1)\right]$$

$$= \sqrt{\frac{N \times (T-3Tlag-5)}{2K \times (T-2K-3m-6)}} \times \left[\frac{(T-3Tlag-3)}{(T-3Tlag-1)} \times W_{N,T}^{Hnc*} - K\right] \rightarrow N(0,1)$$

Additional modifications of Dumitrescu and Hurlin (2012) critical values for fixed N and T samples without and with cross sectional (included in the Matlab code)

In addition, Dumitrescu and Hurlin (2012) also show the critical values for fixed *N* and *T* samples without and with cross sectional. The modified approximated critical values for fixed *N* and *T* samples is

$$\tilde{C}^{*}_{N,T}(\alpha) = Z_{\alpha}\sqrt{N^{-1}var(\widetilde{W}^{*}_{i,T})} + E(\widetilde{W}^{*}_{i,T})$$

$$= Z_{\alpha} \times \frac{(T-3Tlag-1)}{(T-3Tlag-3)} \times \sqrt{\frac{2K}{N} \times \frac{(T-2K-3m-6)}{(T-3Tlag-5)}} + K \times \frac{(T-3Tlag-1)}{(T-3Tlag-3)}$$

and the modified approximated critical values for fixed *N* and *T* samples with cross sectional dependence:

$$\tilde{C}^{*}_{N,T}(\alpha) = Z^{bs}_{\alpha} \times \frac{(T-3Tlag-1)}{(T-3Tlag-3} \times \sqrt{\frac{2K}{N} \times \frac{(T-2K-3m-6)}{(T-3Tlag-5)}} + K \times \frac{(T-3Tlag-1)}{(T-3Tlag-3)}$$