Stock Prices, Exchange Rates and Portfolio Equity Flows: Toda-Yamamoto approach for Granger non-causality test in heterogeneous panels

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Abstract

Portfolio theory predicts that there is a negative correlation between stock prices and exchange rates. This paper draws on transmission channel of foreign portfolio equity investments and tests this by 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). Using macro panel data for eight emerging and developed economies implementing managed or free float exchange rate arrangement, the result shows that Indonesia may be the only case where stock prices affect exchange rates through portfolio equity flows. This study therefore casts doubt on the link between these three variables as expected by the portfolio equity theory.

Keywords: stock prices, exchange rates, portfolio equity flows, portfolio approach,

Granger causality, heterogeneous panels

JEL Classification: F31, G14, G15

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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 then leads to the increased exposure to exchange rate risks. This linkage therefore has important implications for international portfolio management and the impact of stock markets on firm performance.

There are two main competing theories explaining the relationship between stock prices and exchange rates: the traditional approach models (Dornbusch & Fischer 1980) and the portfolio-balance approach models (Frankel 1983).[1] Rather than determined by trade flows as posited by the former, the latter conjectures that the exchange rates are driven by financial market equilibrium conditions. Assuming a home country bias and imperfect substitute between domestic and

[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.

foreign financial assets, Frankel (1983) argues that investors will rebalance their portfolios according to the expected returns of the 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] In short, there is a negative unidirectional causality relationship from stock prices to exchange rates. 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 vice versa. Appreciation or depreciation of the exchange rate will affect both multinational firms (directly) and domestic firms (indirectly). Depending on whether a firm's businesses 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 the current condition of more integrated financial markets, the portfolio-balance approach models seem to receive more empirical support 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)). Interestingly, however, the empirical studies mainly focus on the 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 their neglected essential variable in portfolio-balance models, therefore it is important to include them in order to avoid an omitted variable bias (Granger

[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.

(1969) and Caporale *et al.* (2004)). Hau and Rey (2005) furthermore 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 to exchange rates without incorporating a mediating role for equity portfolio inflows. The portfolio flows are analyzed 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), for example, also mention portfolio equity flows in interpretations of their results, but did not incorporate the 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 concern lack 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 frequency, 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 depicted in Figure 1, can be summarized as follows. Financial liberalisation enables investors to invest their money into 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 in the domestic equity market, either the purchases or sells, will be reflected in the flows of portfolio equity in the balance of payment. The foreign equity flows then will affect exchange rates. With this framework, this paper postulates that a decrease (increase) in stock prices will lead to foreign equity capital outflows (inflows) 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 affect exchange rates.

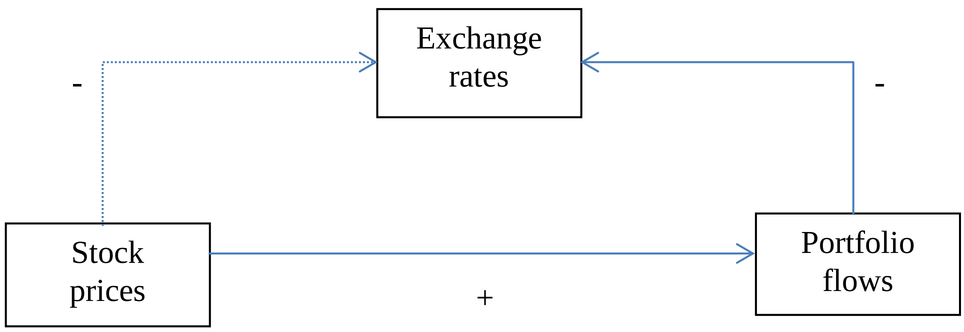

Figure 1. The expected relationship between stock prices, exchange rates and portfolio equity flows

This paper mainly uses a trivariate Granger causality test to examine the relationship among the variables of interests. 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 standard Granger causality test.

The contributions of this paper are (1) integrating portfolio equity inflow variable into a framework with stock price and exchange rate variables, (2) examining this in a panel setting which have a 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 Groenewold and Paterson (2013) which use commodity prices as the intermediating 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 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 a K-th order trivariate panel vector autoregressive as follows:

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

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

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

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 may not only be dependently distributed within a country, but also independently across countries. This paper is interested at testing a Granger (non-)causality between two variables of interest while controlling the other variable.

In such panel VAR, there are at least three different estimation techniques that can be employed: a GMM estimator (Holtz-Eakin *et al.* 1988; Love & Zicchino 2006), a SUR estimator (Konya 2006), and a multivariate least square estimator (Dumitrescu & Hurlin 2012).[3] Except for the first study that is interested at the impulse-response function, the other two propose a different approach to do a Granger causality or non-causality test based on the estimates. 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) offers a causality test for each individual country even though the estimation is conducted with a panel data setting. Konya (2006)'s test can be

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

expanded to trivariate setting but they treat the third variable as an auxiliary variable, not an endogenous one such as in Love and Zicchino (2006).

This paper extends Dumitrescu and Hurlin (2012)'s test to a trivariate setting and relaxes the requirements that the variables of interests must be stationary variables or have the same order of integration. In a context of 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 the asymptotically 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. With this simple alternative approach, the important properties of the panel Wald test statistic proposed by Dumitrescu and Hurlin (2012) still holds. 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 sum all individual country p-values to test the null hypothesis of Granger non-causality. The Fisher test statistic is claimed having a chi-square distribution with $2N$ degrees of freedom when N is fixed and T goes to infinity.

For a bivariate setting with both variables Y_i and X_{1i} are stationary,

$$y_{i,t} = \alpha_i + \sum_{p=1}^{K} \beta_{i,p}\, y_{i,t-p} + \sum_{p=1}^{K} \gamma_{i,p}\, X_{1i,t-p} + \varepsilon_{i,t} \tag{4}$$

With the null and alternative hypotheses:

$H_0: \gamma_i = 0 \quad \forall i = 1, \dots, N$

$H_1: \gamma_i = 0 \quad \forall i = 1, \dots, N_1; \ \gamma_i \neq 0 \quad \forall i = N_1 + 1, \dots, N$

Dumitrescu and Hurlin (2012) show that a panel Wald test statistic

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

will be asymptotically distributed according to a normal distribution with mean zero and variance equals to one as $T \to \infty$, where $W_{N,T}^{Hnc} = \frac{1}{N}\sum_{i=1}^{N} W_{i,T}$ and $W_{i,T}$ is the individual Wald test statistic for i-th cross-section unit corresponding to the individual test $H_0: \gamma_i = 0$. For a fixed dimension of T to make that normal distribution still holds,[4] however, the panel statistic needs to be standardized and modified to $\tilde{Z}_{N,T}^{Hnc}$ as follows:

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

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

$$y_{i,t} \quad = \alpha_i + \sum_{p=1}^{K+m} \beta_{i,p}\, y_{i,t-p} + \sum_{p=1}^{K+m} \gamma_{i,p}\, X_{1i,t-p} + \sum_{p=1}^{K+m} \delta_{i,p}\, X_{2i,t-p} + \varepsilon_{i,t} \tag{7}$$

Let $Y_i = [y_{i,1} \colon y_{i,2} \colon \dots \colon y_{i,K+m}]'$, $X_{1i} = [x_{1i,1} \colon x_{1i,2} \colon \dots \colon x_{1i,K+m}]'$ and $X_{2i,t} = [x_{2i,1} \colon x_{2i,2} \colon \dots \colon x_{2i,K+m}]'$ are $(T, K+m)$ matrices, respectively. All three variables are endogenous with the maximum order of integration m and $X_{2i,t}$ is

4 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 performed well when N is small as in our case.

hold 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; and $\hat{\theta}_i$ and $\hat{\varepsilon}_i$ are the OLS estimator for θ_i and the residuals from the regression model (4), respectively. 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:[5]

$$W_{i,T}^{*} = \hat{\theta}_i' R' [\hat{\sigma}_i^2 R (Z_i' Z_i)^{-1} R']^{-1} R \hat{\theta}_i = \frac{\hat{\theta}_i' R' [R (Z_i' Z_i)^{-1} R']^{-1} R \hat{\theta}_i}{\hat{\varepsilon}_i' \hat{\varepsilon}_i / (T - 3Tlag - 1)} \quad (7)$$

$$Z_{N,T}^{Hnc*} = \sqrt{\frac{N}{2K}} (W_{N,T}^{Hnc} - K) \quad (8)$$

$$\tilde{Z}_{N}^{Hnc*} = \sqrt{\frac{N \times (T - 3Tlag - 5)}{2K \times (T - 2K - 3m - 6)}} \times \left[\frac{(T - 3Tlag - 3)}{(T - 3Tlag - 1)} \times W_{N,T}^{Hnc} - K\right] \quad (9)$$

To accommodate cross-sectionally dependence, Dumitrescu and Hurlin (2012) propose to use bootstrapped critical values. Steps to modify their bootstrapping technique in a trivariate Toda and Yamamoto (1995)'s approach is as follows:[6]

1. Estimate model (7) under the null hypothesis, that is set $\gamma_{i,p} = 0, \forall p = 1, \ldots, K$ 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;

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[5] See Appendix I for proofs.

[6] A Matlab code built based on the codes written by Hurlin (http://www.runmycode.org/companion/view/42) and by Emirmahmutoglu (http://www.runmycode.org/companion/view/89) is used to run the trivariate Granger non-causality tests. The code is available upon request.

3. Construct a resample 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 construct the 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 (*EqFlows*) are then calculated by subtracting assets from liabilities of the net portfolio investments and expressed as the percentage of current GDP. The MSCI series for the end of period exchange rates per US dollar (*Currency*) and the stock indices (*Index*) are collected from Thomson Reuters Datastream Professional.[7]

The dataset comprises eight economies covering both advanced and emerging markets implementing managed or free float exchange rate arrangement i.e. Australia, Canada, Indonesia, Japan, South Korea, Sweden, Thailand and U.K. The sample period of 1993:Q1-2008:Q4 are mainly chosen to make sure the completeness of series having the longest available quarterly measures. Our

[7] It is interesting to notice 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.

sample therefore consists of eight cross section units and 64 time series units (N = 8 and T = 64). All series except the net portfolio capital inflows are in a natural logarithm.

This paper examines both an individual time series and a panel data to get a more comprehensive result. For the individual time-series, this paper employs the unit root tests of Zivot and Andrews (2002) and the cointegration test of Gregory and Hansen (1996). These tests allow for the presence of structural breaks in the time series. For the panel data, this paper employs the diagnostic tests for cross section dependence of Pesaran (2004) and the modified Sargan-Bhargava (MSB) panel unit root test of Bai and Carrion-i-Silvestre (2009) to test if our series are cross-sectionally dependent and whether they contain unit roots, respectively. 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 exists. To assess whether the series are co-integrated, the panel cointegration test with structural breaks and cross-section dependence of Banerjee and Carrion-i-Silvestre (2013) which also use common factors to model the cross-sectional dependence are employed.

4. Results

4.1. Descriptive Statistics

The individual country time-series plots for all variables of interest are provided in Figure A1-A3 in Appendix II. An increasing trend in stock prices and exchange rates is a common feature in all economies; while the variable of net capital inflows of portfolio equity seems more fluctuate around 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.

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
Std.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***

Note: 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 ***).

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 equity inflows. Table 1 also indicates the presence of cross-country dependence among economies in our sample, which may be due to the financial market integration or spillover 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. The tests show in general that $EqFlows_{i,t}$ is I(0), while $Index_{i,t}$ and $Currency_{i,t}$ are I(1). For Indonesian Rupiah and Thailand Baht, however, 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. 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 cointegration relationship may also exist for Indonesia.

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***

Note: The null hypothesis assumes that all series are non-stationary. The statistics are computed for the model allowing to have 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 and 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

Note: 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 to have 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 and 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 shows that the maximum lag length to be used in a standard VAR model may vary depending on the criteria used. However in general 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 the maximum lag length varies from one to two. Only Indonesia is indicated to have the maximum number of lag of four. Based on these criteria, we decide that the maximum lag is either one ($K = 1$) or two ($K = 2$).

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

Note: 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.

The results for Granger causality test in 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 least 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 existence cointegration relationship between the two variables in Table 3 also confirms these causality test results, at least for Korea and Thailand. However, column (4) provides 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.

TABLE 5
Trivariate Granger causality tests using Toda Yamamoto Framework

	$Index_{i,t}$ → $Currency_{i,t}$	$Index_{i,t}$ → $EqFlows_{i,t}$	$EqFlows_{i,t}$ → $Index_{i,t}$	$EqFlows_{i,t}$ → $Currency_{i,t}$	$Currency_{i,t}$ → $Index_{i,t}$	$Currency_{i,t}$ → $EqFlows_{i,t}$
	(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
	−	+	+	+	−	+

Note: → 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

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 approach, in particular the risk rebalancing channel for portfolio equity flow approach, is only supported in case of Indonesia.

Compared to the results of bivariate analysis done by other studies, this trivariate study provides some similar findings. No evidence for Granger causality between stock prices and exchange rates in Australia is similar to the findings of Hau and Rey (2005) and Groenewold and Paterson (2013). Similarly, no evidence for Granger causality in Japan is also implied by the studies of Granger *et al.* (2000), Hau and Rey (2005), and Caporale *et al.* (2014). A unidirectional causality from stock prices to exchanges rates in UK found in Hau and Rey (2005) and Caporale *et al.* (2014) is also the case in this study. In case of Korea and Thailand this study supports the feedback relations as found in 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 two variables, this study fails to find such relation in Canada and Sweden. In particular for Indonesia, this study's finding may resolve conflicting findings from other studies. 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).

5.2. 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 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 then 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).

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***

Note: 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***

Note: 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.

The next step is to examine the possibility of cointegration relationship between our I(1) series: $Currency_{i,t}$ and $Index_{i,t}$. As an alternative for Banerjee and Carrion-i-Silvestre (2013) test, this paper 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 more preferable than the other panel cointegration tests. Table 8 summarizes the three panel cointegration tests and provides insufficient evidence to reject the null hypothesis of no-cointegration.

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.[8] The results of the trivariate Toda-Yamamoto approach for Granger non-causality test in heterogeneous panels are summarized in Table 9 below.

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-balanced approach.

[8] 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.

TABLE 8
The Panel Cointegration tests of Westerlund (2007), Banerjee and Carrion-i-Silvestre (2013) and Di Iorio and Fachin (2014)

Test	Statistic	Critical value/p-value[1]
Banerjee and Carrion-i-Silvestre (2013)[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

Note:
[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 number of lags, and the Bartlett kernel window width set of 4. The p-values are robust to cross sectional dependence and computed with 500 number of bootstrap replications.
[4] ADF statistics are computed for the model with constant and trend, maximum two number of 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

Note: → 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.

5. Conclusion

In this paper, we have examined the portfolio approach postulating the negative causal relationship between stock prices and exchange rates through portfolio capital flow transmission channel. To this end, this paper has developed a trivariate Granger non-causality test in heterogeneous panels to a trivariate setting in a framework of Toda and Yamamoto's approach for possible non-stationary variables. Using macro panel data setting for eight emerging and developed economies with managed or fee floating exchange rate arrangement, this study casts doubt on the linking among the three variables. Indonesia may be the only case where stock prices affect exchange rates through portfolio equity flow channel.

APPENDIX I.

A modified Dumitrescu and Hurlin (2012)'s individual Wald test can be calculated as: $W_{i,T}^{*}$ is essentially similar to $W_{i,T}$ with different definitions of matrices $\hat{\theta}_i', R, Z_i$ $\quad W_{i,T}^{*} = \hat{\theta}_i' R' [\hat{\sigma}_i^2 R (Z_i' Z_i)^{-1} R']^{-1} R \hat{\theta}_i = \frac{\hat{\theta}_i' R' [R(Z_i' Z_i)^{-1} R']^{-1} R \hat{\theta}_i}{\hat{\varepsilon}_i' \hat{\varepsilon}_i / (T - 3Tlag - 1)}$ and when $T \to \infty$, $W_{i,T}^{*} \xrightarrow{d} \chi^2(k), \forall_i = 1, \dots, N$. Furthermore, also when $N \to \infty$, $E(W_{i,T}^{*}) = K$ and $Var(W_{i,T}^{*}) = 2K$, 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 $Z_{N,T}^{Hnc*} \to N(0,1)$.

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

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

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

Proof:

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

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

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

The denominator can be simplified as follows:

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

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

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

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

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

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

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

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,

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

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

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

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

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

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

In addition, Dumitrescu and Hurlin (2012) also show the critical values for 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:

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

APPENDIX II.

Line Plots of Portfolio Equity Flows, Exchange Rates and Stock Prices

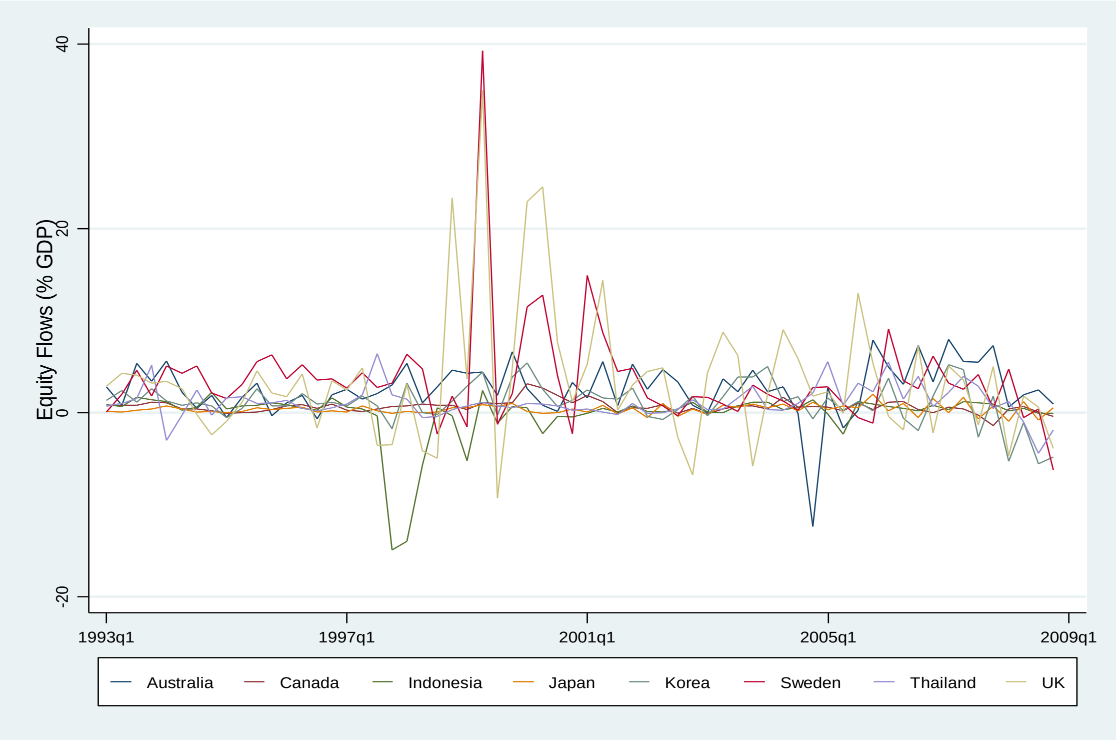

Figure A1. Portfolio Equity Flows (as percentage of GDP)

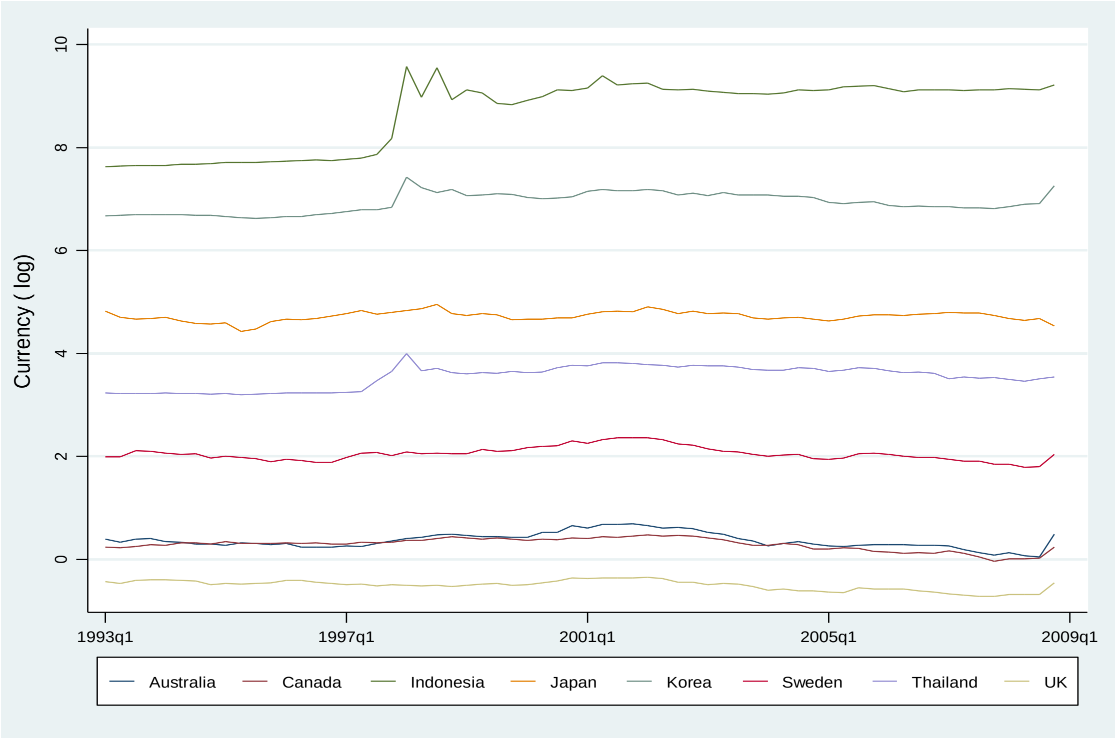

Figure A2. Exchange Rates (in natural logarithm)

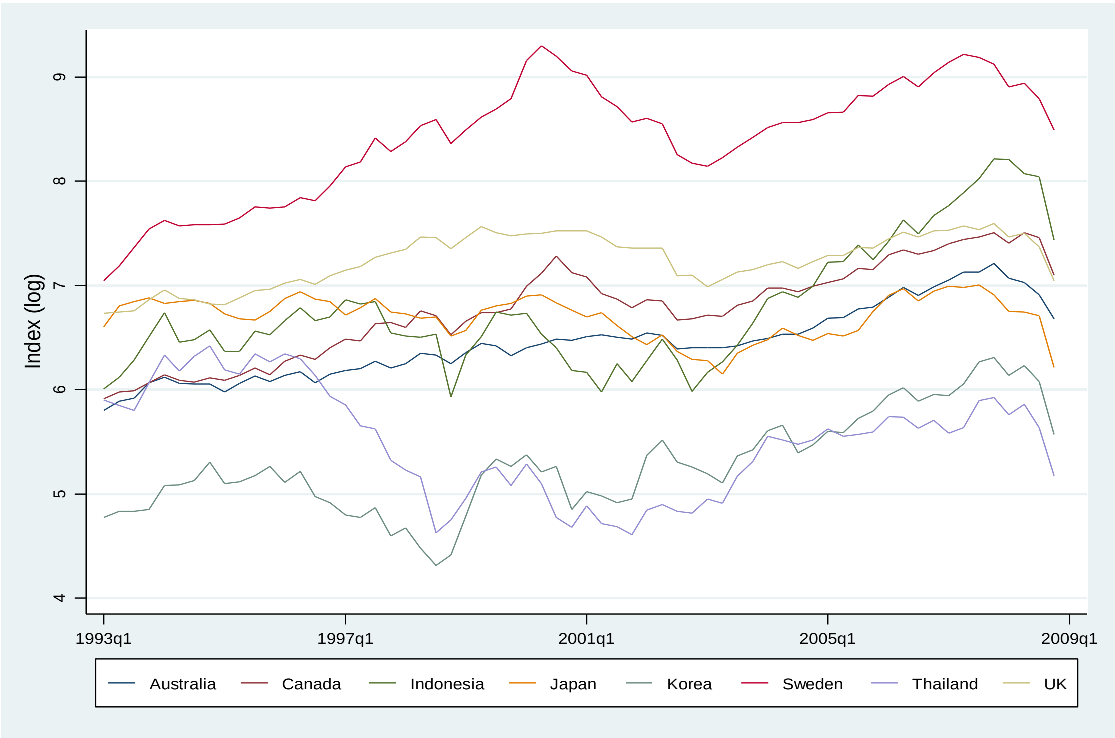

Figure A3. Stock Prices (in natural logarithm)

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