Economic Modelling 78 (2019) 209–224

Contents lists available at ScienceDirect

Economic Modelling

journal homepage: www.journals.elsevier.com/economic-modelling

Exchange rates and fundamentals: A bootstrap panel data analysis

Zixiong Xie[a], Shyh-Wei Chen[b,*]

[a] Department of International Economics and Trade, Institute of Resources, Environment, and Sustainable Development, College of Economics, Jinan University, Guangzhou 510632, China

[b] Department of International Business, Tunghai University, Taiwan, ROC

ARTICLE INFO

JEL classification:
C32
Q43

Keywords:
Exchange rate
Fundamental
Bootstrap
Panel model
Granger non-causality

ABSTRACT

This study attempts to re-examine the Granger non-causality from exchange rates to observed fundamentals based on the present value model of Engel and West (2005). To this end, we employ the bootstrap panel Granger non-causality analysis, which allows us to untangle the causal nexus between exchange rates and fundamentals in panel data. Among the main results, it is found that the null hypothesis of no cross-sectional dependence across the members of the panel is strongly rejected, indicating that the bootstrap critical value is required in conducting the panel Granger non-causality test. The null hypothesis of Granger non-causality running from the fundamentals to exchange rates is significantly rejected, implying that the monetary approach of exchange rate determination is a useful benchmark to understand the evolution of the exchange rate. Empirical evidences also show that exchange rates Granger-case the fundamentals, supporting the view that exchange rates are determined as the present value that depends in part on observed fundamentals.

1. Introduction

The monetary model of exchange rate plays an important role of exchange rate determination in international finance. There are two well-known versions of the monetary model. The first one is the 'Chicago' theory or the flexible price monetary model introduced by Frenkel (1976), Mussa (1976) and Bilson (1978a, 1978b). The second is the 'Keynesian' or the sticky price monetary model proposed by Dornbusch (1976) and Frankel (1979). Theoretically, the monetary model states that in the long run, exchange rates are determined by a number of economic fundamentals such as money supplies, real outputs, interest rates and inflation rates. However, the empirical studies (e.g., Sarantis, 1994; Cushman, 2000; Engel and West, 2005) do not provide solid and consistent evidence in support of such relationship. In particular, the seminal work of Meese and Rogoff (1983a, 1983b) shows that the monetary model cannot outperform a simple random walk model in terms of the out-of-sample forecasts.[1] This delicate relationship between exchange rates and economic fundamentals, which is termed the "exchange rate disconnect puzzle" and coined by Obstfeld and Rogoff (2001), becomes a stylized fact documented in international macroeconomics (Mark, 1995; Cheung et al., 2005; Sarno, 2005).[2]

A myriad of studies have devoted a great deal of effort to dissect this puzzle, i.e., to provide empirical evidence in support of the monetary model of exchange rate determination. Many studies use the cointegration approach to examine the long-run relationship between exchange rates and market fundamentals (for example, to name a few, Sarantis, 1994; Pilbeam, 1995; Mark, 1995; Cushman, 2000; Francis et al., 2001; Rapach and Wohar, 2002; Frenkel and Koske, 2004; Civcir, 2004;

* Corresponding author. Department of International Business, Tunghai University, No.1727, Sec.4, Taiwan Boulevard, Xitun District, Taichung 40704, Taiwan, ROC.

E-mail addresses: txiezixiong@jnu.edu.cn (Z. Xie), schen@thu.edu.tw (S.-W. Chen).

[1] Wu and Wang (2013) survey the literature and point out that there are three reasons to explain the failure of defeating the random walk benchmark. The first one refers to the imprecise parameter estimates. The second reason refers to measurement errors in fundamental exchange rates. The third reason refers to the failure of including useful information from other relevant variables. Baxter and Stockman (1989) and Flood and Rose (1995) also provide an explanation why monetary model alone cannot explain the high variability of the exchange rates. This is because the transition from fixed to floating exchange rates leads to a strong increase in nominal and real exchange rate variability that is not followed by a similar increase in the variability of macroeconomic fundamentals.

[2] See Neely and Sarno (2002), Sarno (2005) and Laganá and Sgro (2007) for selective overviews of the exchange rate disconnect puzzle. Evans (2010) has recently proposed a theoretical model of exchange-rate determination that bridges the gap between existing microstructure and traditional models. It provides a straightforward solution to the exchange-rate disconnect puzzle. Namely, the high frequency behavior of spot exchange rates reflects the flow of new information reaching dealers concerning the slowly evolving state of the macroeconomy, rather than the effects of shocks that drive rapidly changing macroeconomic conditions.

https://doi.org/10.1016/j.econmod.2018.09.021
Received 20 April 2018; Received in revised form 12 September 2018; Accepted 19 September 2018
Available online 24 September 2018

Islam and Hansan, 2006; Chang and Su, 2014; Bahmani-Oskooee et al., 2015; Tawadros, 2001, 2017).[3] Motivated by the statistical power of the advances in panel unit root and panel cointegration tests (Maddala and Wu, 1999; Westerlund, 2006, 2014, 2015), an increasing number of authors (for example, Groen, 2000, 2002; Mark and Sul, 2001; Rapach and Wohar, 2004; Basher and Westerlund, 2009; Cerra and Saxena, 2010; Beckmann et al., 2012; Dabrowski et al., 2014) have applied these new tools to test the monetary model of exchange rate in the long run.[4] Basically, an important feature of previous studies is that distinct results based on previous research are due to differences in methodology, approaches and samples and are subject to diverse interpretations, thus making it difficult to reach a corroborative position on the puzzle. Cheung et al. (2005) emphasize that no particular model is dominated and conclude that it may be that one model will do well for one exchange rate, and not for another.

Engel and West (2005), recently, demonstrate that the frail link between nominal exchange rates and economic fundamentals can be reconciled within a rational expectation model. Based on Campbell and Shiller (1988a, 1988b), they prove that exchange rates are the present discounted value of expected economic fundamentals and their stochastic behavior look like a random walk. This is because nonstationary fundamentals will impart nonstationarity to exchange rates. Moreover, a large discount factor gives greater weight to expectations of future fundamentals relative to current fundamentals. As a result, current fundamentals are only weakly related to exchange rates as exchange rates appear to follow an approximate random walk (Balke et al., 2013). As such, Engel and West's (2005) result hinges critically upon two assumptions: (i) economic fundamentals are nonstationary (or near-random walk) processes, and (ii) the factor for discounting expected economic fundamentals in the exchange rate equation is relatively large, e.g., smaller than unity but greater than 0.9.[5]

From an economic point of view, the seminal work of Engel and West (2005) provides an important theoretical foundation to explain the exchange rate disconnect puzzle. That is, it might well be that exchange rates are determined by economic fundamental variables, but in many occasions the exchange rates are in fact well approximated as random walks. Their study attracts a great deal of attention from the other researchers. For example, Balke et al. (2013) examine the degree to which fundamentals can explain the fluctuations of exchange rates by using the state-space model to decompose the contribution of observed and unobserved shocks. Bekiros (2014) employs the nonlinear Granger non-causality test, proposed by Diks and Panchenko (2005, 2006), to examine the causal relationships between the exchange rates and economic fundamentals. Chen and Chou (2015) reveal the fluctuations in exchange rates and fundamentals using the permanent-transitory decomposition method developed by Gonzalo and Ng (2001).

From an applied economic point of view, an important empirical implication of the Engel and West (2005) present value model is that current exchange rates are helpful for predicting the future fundamentals in terms of Granger's definition.[6] This is because much of the short-run fluctuation in exchange rates is driven by changes in expectations about the future. Assuming that the present value model is a good approximation and that expectations reflect information about future fundamentals, changes of exchange rates will be useful in forecasting these fundamentals. By using the Johansen (1991) cointegration approach for the data of G7 nations, Engel and West (2005) find almost no evidence of cointegration in the multivariate case, and only 5 out of 24 cases in the bivariate cases and find that there are significant causality relations from exchange rates to fundamentals. Ko and Ogaki (2015) and Bahmani-Oskooee et al. (2015) use the exact same dataset of Engel and West (2005) to re-examine the causal relationship between exchange rates and economic fundamentals. Ko and Ogaki (2015) claim that the findings of Engel and West (2005) might suffer from the small-sample problem. Therefore, Ko and Ogaki (2015) suggest to use the bootstrap method to re-evaluate the Granger non-causality from exchange rates to fundamentals with the bivariate vector autoregressive (VAR) model since no evidence of cointegration is found between exchange rates and economic fundamentals using Johansen's method. Their bootstrap test results show that the Granger causality from exchange rates to the observable fundamentals is not as significant as the existing evidence based on the asymptotic distribution in all sample periods. Bahmani-Oskooee et al. (2015) emphasize the disadvantages of the Johansen cointegration approach and instead adopt the autoregressive distributed lag (hereafter ARDL) approach to cointegration, which is proposed by Pesaran et al. (2001), to implement the empirical study.[7] On the contrary, they report cointegration in all 6 multivariate cases, and 20 out of 24 bivariate cases. Furthermore, their Granger non-causality tests report evidence that macroeconomic fundamentals help predict exchange rates in both the short run and long run. However, their empirical results provide weak evidence in support of Granger causality running from exchange rates to macroeconomic fundamentals.

The aim of this study is to revisit the Granger causal relationships between exchange rates and observed fundamentals in the context of the monetary model. To this end, we decide to adopt the panel data analysis. The reason to employ the panel data approach is motivated by the following studies. The first comes from the seminal work of Engel and West (2005). On page 512 of their paper, they suggest to employ the panel data approach in the future because "the empirical results are not uniformly strong. As well, it remains to be seen how well they hold upon, for example, use of panel data …". The second is motivated by the studies of Bahmani-Oskooee et al. (2015) and Ko and Ogaki (2015). The former does find evidence of cointegration between exchange rates and economic fundamentals, but the latter does not. The frailty of Bahmani-Oskooee et al. (2015) is that it suffers from the pre-test bias. Likewise, an inefficiency of Ko and Ogaki's (2015) study is that they implicitly assume that the variables are required to be covariance stationary. Hence, they use the first-difference of the data to conduct the Granger non-causality analysis since exchange rates and economic fundamentals are integrated of order one. However, using the first-difference of the data can capture the short-run dynamic causal relations among variables, but it neglects the long-run equilibrium relationship if there exist

[3] Bahmani-Oskooee et al. (2010) stipulate that (i) evidence of cointegration between exchange rates and market fundamentals and (ii) parameters showing the proper sign and significance in the long-run relationship are necessary conditions in support of the monetary model of exchange rate determination.

[4] Another avenue to examine this issue is adopting the nonlinear approach. Many studies have supported nonlinear, mean-reverting adjustment of real exchange rates and has shown that the exponential smooth transition autoregressive (ESTAR) model provides a parsimonious fit to the data. See, for example, Michael et al. (1997), Taylor and Peel (2000), Taylor et al. (2001), Kilian and Taylor (2003), Clarida et al. (2003), Wu and Hu (2009) and Kim et al. (2010) for details. Alternative, Frömmel et al. (2005a, 2005b), Grauwe and Vansteenkiste (2007) and Yuan (2011) examine the monetary exchange rate model using the Markov switching model. Engel and Hamilton (1990), Engel (1994), Shen and Chen (2004), Klaassen (2005) and Yuan (2011) also try to evaluate the forecasting performance of exchange rates by employing the Markov switching model. Beckmann et al. (2011) investigate the temporal stability of the relationship between the Deutschmark/US dollar exchange rate and macroeconomic fundamentals by applying a time-varying coefficient approach.

[5] The first assumption of nonstationary fundamentals has been supported by many studies such as Engel and West (2005), Engel et al. (2007), Bahmani-Oskooee et al. (2015), Ko and Ogaki (2015) and Tawadros (2017). While the second assumption of a large discount factor has been sustained by Engel and West (2004) and Sarno and Sojli (2009).

[6] Variable s_t is said to Granger-cause variable f_t if f_t can be better predicted using the histories of both s_t and f_t than it can by using the history of f_t alone.

[7] Panopoulou and Pittis (2004) show that in finite sample the ARDL approach is superior to the other cointegration approaches, including Johansen's method, due to the fact other approaches suffer from 'truncation bias'.

a cointegration relationship between empirical variables and, therefore, the results of Granger non-causality test could be misleading. Ko and Ogaki (2015) find that exchange rates do not help predict fundamentals because the null hypothesis of Granger non-causality running from exchange rates to fundamentals is not rejected. However, they conclude that "we do not intend to conclude that the present-value model under Engel and West's explanation for the exchange rate is wrong, …In fact, even after extending the data span as what was done in Engel et al. (2007), the evidence for the causality from the exchange rate to the observable fundamental is also not uniformly strong. Hence, studies that help to explore the causality relation from exchange rates to fundamentals could be a priority for the future research (Ko and Ogaki, 2015, p. 205)."

In this study, we employ Emirmahmutoglu and Kose's (2011) panel Granger non-causality test to avoid the pre-test bias and thereby mitigate the problem of using the first-difference of the data. As already outlined, most previous studies (Bahmani-Oskooee et al., 2015; Ko and Ogaki, 2015; Tawadros, 2017) adopt linear Granger non-causality test and cointegration approaches to examine the causal relationship between exchange rates and macroeconomic fundamentals. However, it is well-known that the standard asymptotic theory is not applicable to hypothesis testing in the level VAR model if the variables are integrated or cointegrated. This is because the usual Wald test statistics for Granger non-causality based on the level VAR model are not only characterized by a non-standard asymptotic distribution, but depend on nuisance parameters in general if the variables are non-stationary (Toda and Phillips, 1993). Therefore, a pre-test is needed to determine the order of integration of variables before estimating the appropriate VAR model from which statistical inferences are derived. However, the Granger non-causality test may suffer from severe pre-test bias. The methodology proposed by Emirmahmutoglu and Kose (2011) can avoid this bias because it is an extension of Toda and Yamamoto's (1995) test. Toda and Yamamoto (1995) recommend using a modified Wald (MWALD) test in a lag augmented vector autoregression (LA-VAR) which has a conventional asymptotic chi-square distribution when VAR($p + dmax$) is estimated, where p is the lag order and $dmax$ is the maximal order of integration suspected to occur in the process. The only prior information needed for the LA-VAR approach is the maximum order of integration of the processes.

In light of the fact that the pre-tests for a unit root and cointegrating rank (or taking differences in the data) are not required, the associated pre-test bias and size distortion can be avoided, at least asymptotically (Yamada and Toda, 1998). The simulation study by Emirmahmutoglu and Kose (2011) shows that their test has good power and reasonable size performances even if N and T are small. To the best of the authors' knowledge, no other study in the literature has ever examined the causal relationships between exchange rates and economic fundamentals by using Emirmahmutoglu and Kose's (2011) approach.[8] It thus allows us to untangle the causal relationship between exchange rates and macroeconomic fundamentals and helps us to discriminate between competing theories.

The remainder of this paper is organized as follows. Section 2 reviews the monetary approach of the exchange rate determination and the present value model of exchange rate. Section 3 briefly discusses the model specification of exchange rates and macroeconomic fundamentals and introduces the econometric methodology that we employ, and Section 4 describes the data and the empirical test results. Section 5 presents the conclusions that we draw from this research.

[8] Kónya (2006) suggests a different panel non-causality test which is based on the Seemingly Unrelated Regressions (SUR) estimator proposed by Zellner (1962), and the Wald test with country-specific bootstrap critical values. Kónya's (2006) test does not require pretesting for unit roots and cointegration apart from the lag structure. Nonetheless, this is an important problem since the unit-root and the cointegration tests in general suffer from low power and different tests often lead to contradictory results.

2. Review of the monetary and present-value models of exchange rate determination

2.1. The flexible price monetary model

The flexible price monetary model, championed by Frenkel (1976) and Bilson (1978a, 1978b),[9] is based on three assumptions: the purchasing power parity (hereafter PPP), money market equilibrium and the uncovered interest parity (hereafter UIP). It assumes that goods prices are perfectly flexible and thus that purchasing power parity holds instantaneously:

$$s_t = p_t - p_t^*, \tag{1}$$

where s_t is the log of the spot exchange rate, defined as the price of foreign currency in terms of domestic and p_t and p_t^* are logs of the domestic and foreign price levels, respectively. The quantity theory of money posits that prices are determined by equilibrium in money market. That is, we assume a simple money demand functions at home and abroad,

$$m_t = p_t + \phi y_t - \lambda i_t, \tag{2}$$

$$m_t^* = p_t^* + \phi y_t^* - \lambda i_t^*, \tag{3}$$

where m_t and m_t^* are the logs of the domestic and foreign money supplies, respectively; y_t and y_t^* are the logs of domestic and foreign real income; and i_t and i_t^* the domestic and foreign interest rate. For simplicity, the elasticity with respective to income, ϕ, and the semielasticity with respective to the interest rate, λ, are assumed to be equal across countries. Combining Eqs. (1)–(3) yields one representation of the flexible price monetary equation (Bilson, 1978a, 1978b):

$$s_t = (m_t - m_t^*) - \phi(y_t - y_t^*) + \lambda(i_t - i_t^*). \tag{4}$$

If it is to maintain that bond supplies do not affect interest rate or exchange rates money supplies do, the monetary approach must assume that domestic and foreign bonds are perfect substitutes and thus that uncovered interest parity holds,

$$\Delta s_t^e = E_t s_{t+1} - s_t = i_t - i_t^*, \tag{5}$$

where Δs^e is the expected depreciation of domestic currency. The market will be aware of the PPP condition and so we will have

$$\Delta s_t^e = \pi_t - \pi_t^*, \tag{6}$$

where π_t and π_t^* are the expected inflation rates, at home and abroad, respectively. Substituting Eqs. (5) and (6) into Eq. (4), we have an alternative representation of the flexible price monetary model (Frenkel, 1976):

$$s_t = (m_t - m_t^*) - \phi(y_t - y_t^*) + \lambda(\pi_t - \pi_t^*). \tag{7}$$

Eqs. (4) and (7) stipulate that exchange rates, as the relative price of money, are determined by the supply and demand for money. An increase in the domestic money supply (relative to the foreign money supply) causes a proportionate depreciation of the domestic currency. An increase in the domestic money demand, such as results from an increase in domestic income or a decrease in the domestic interest rate, causes an appreciation of domestic currency. Since goods prices are flexible, a rise in the domestic relative to the foreign interest rate reflects a corresponding rise in the domestic relative to the foreign inflation rates. As such, an increase in domestic interest rates leads to a fall in the demand for the domestic relative to the foreign currency, causing it to depreciate. Thus, there is a *positive* relationship between exchange rates and the interest rates differential.

[9] The contents of Sections 2.1 and 2.2 draw heavily from Frankel (1984).

2.2. The sticky price monetary model

The alternative class of monetary models to exchange rate determination, proposed by Dornbusch (1976) and Frankel (1979), posits that good prices are rigid, and thus purchasing power parity is assumed to hold in the long run:

$$\bar{s} = \bar{p} - \bar{p}^*, \tag{8}$$

where a 'bar' over a variable denotes long-run equilibrium. This Eq. (7) holds only in long-run equilibrium:

$$\bar{s} = (\overline{m} - \overline{m}^*) - \phi(\bar{y} - \bar{y}^*) + \lambda(\bar{\pi} - \bar{\pi}^*). \tag{9}$$

In the short run, the spot rate can deviate from its equilibrium value, but the market expects the spot rate to regress toward equilibrium at a rate proportional to the gap:

$$\Delta s_t^e = -\theta(s_t - \bar{s}) + \bar{\pi} - \bar{\pi}^*. \tag{10}$$

This form of expectations turn out to be rational in a model in which prices adjust gradually over time in response to excess goods demand but also move in line with the underlying inflation rate $\bar{\pi}$. Combining (10) with the monetary approach's assumption of uncovered interest rate parity (5), which is retained in the Dornbusch model, we have an expression for the gap between the current spot rate and its equilibrium level:

$$s_t - \bar{s} = -\frac{1}{\theta}[(i_t - \bar{\pi}) - (i_t^* - \bar{\pi}^*)]. \tag{11}$$

A tight monetary policy raises the real interest rate differential, attracts a capital inflow, and appreciates the currency above its equilibrium value.

By combining Eqs. (9) and (11), we obtain the sticky price monetary equation of exchange rate determination:

$$s_t = (\overline{m} - \overline{m}^*) - \phi(\bar{y} - \bar{y}^*) - \frac{1}{\theta}(i_t - i_t^*) + \left(\lambda + \frac{1}{\theta}\right)(\bar{\pi} - \bar{\pi}^*). \tag{12}$$

Adding an error term and together with the assumption that the long-run values of money and real income are given by their current values, we have the Frankel (1979) real interest differential model:

$$s_t = \alpha + \beta(m_t - m_t^*) - \phi(y_t - y_t^*) + \gamma(i_t - i_t^*) + \eta(\pi_t - \pi_t^*) + \varepsilon_t. \tag{13}$$

where $\beta = 1$, $\gamma = -\frac{1}{\theta}$, $\eta = \left(\lambda + \frac{1}{\theta}\right)$, and η is assumed to greater than γ in absolute value ($\eta > \gamma$). The Dornbusch (1976) sticky-price monetary model is nested within Eq. (13) with the restriction that the coefficient on the expected inflation differential is equal to zero ($\eta = 0$). Since goods prices are sticky, it posits that a rise in the domestic relative to the foreign interest rate reflects a drop in domestic money supply. Prices in the goods market will not adjust instantaneously, but will gradually fall. As such, an increase in domestic interest rates relative to its foreign counterpart will cause capital inflow, causing an appreciation in the domestic currency. Thus, there is a *negative* relationship between exchange rates and the interest rates differential, i.e., $\gamma < 0$. For readers' information, we summarize the predicted sign for the four variants of the monetary model of the exchange rate determination in Table 1.

2.3. The present-value model of exchange rate

Engel and West (2005) demonstrate that the nominal exchange rates is the present discounted value of the expected future economic fundamentals and it is applicable to many exchange rate determination models. We use the flexible price monetary model as an example. First, since the purchasing power parity in general only holds in the long run (Rogoff, 1996), we modify the PPP in Eq. (1) as follow:

$$s_t = p_t - p_t^* + v_t^{ppp}, \tag{14}$$

where the variable v_t^{ppp} is to pick up those deviations from PPP.

Next, we add the unobserved money demand disturbance to the classical monetary model as below:

$$m_t = p_t + \phi y_t - \lambda i_t + v_t^{md}, \tag{15}$$

$$m_t^* = p_t^* + \phi y_t^* - \lambda i_t^* + v_t^{*md}, \tag{16}$$

where v_t^{md} and v_t^{*md} represent unobserved variables that shift money demand. Combining Eqs. (14)–(16) yields the following relationship for the nominal exchange rates:

$$s_t = (m_t - m_t^*) - \phi(y_t - y_t^*) + \lambda(i_t - i_t^*) - (v_t^{md} - v_t^{*md}) + v_t^{ppp}. \tag{17}$$

Finally, following Balke et al. (2013), we integrate the model a generalized uncovered interest parity (UIP) condition that allows for a time-varying risk premium, v_t^{uip}, since the in general UIP does not sustain and the equilibrium model would imply a non-trivial risk premium (Engel, 1996):

$$i_t - i_t^* = E_t s_{t+1} - s_t + v_t^{uip}, \tag{18}$$

where $E_t s_{t+1} - s_t$ is the expected depreciation of domestic currency. Substituting Eq. (18) into Eq. (17), we can derive a stochastic difference equation that describes how the nominal exchange rates would depend upon observed monetary fundamentals and unobserved disturbances.

$$s_t = \frac{1}{1+\lambda}(m_t - m_t^*) - \frac{\phi}{1+\lambda}(y_t - y_t^*) + \frac{1}{1+\lambda}(v_t^{ppp} - v_t^{md} + v_t^{*md}) + \frac{\lambda}{1+\lambda}v_t^{uip} + \frac{\lambda}{1+\lambda}E_t s_{t+1}. \tag{19}$$

Eq. (19) can be manipulated so as to express the exchange change rate in terms of its deviation from observed fundamentals, similar to the stock price decomposition by Campbell and Shiller (1988a, 1988b):

$$s_t - f_t = \psi \cdot E_t[s_{t+1} - f_{t+1}] + \psi \cdot E_t[\Delta f_{t+1}] + R_t. \tag{20}$$

where $f_t \equiv (m_t - m_t^*) - \phi(y_t - y_t^*)$ is the observed monetary fundamental and $s_t - f_t$ is the deviation of the current exchange rates from their current observed monetary fundamentals. Through this paper, we set $\phi = 1$ as in Mark (1995) and Rapach and Wohar (2002). Term $\psi = \frac{\lambda}{1+\lambda}$ is the so-called discount factor. The unobserved term, R_t, consists of the unobserved money demand shifter as well as deviations from both uncovered interest rate parity and purchasing power parity:

$$R_t = \frac{1}{1+\lambda}(v_t^{ppp} - v_t^{md} + v_t^{*md}) + \frac{\lambda}{1+\lambda}v_t^{uip}. \tag{21}$$

Iterating Eq. (20) forward and under the assumption of no bubble condition the model can be solved as follows:

$$s_t - f_t = E_t\left[\sum_{j=1}^{\infty}\psi^j \Delta f_{t+j}\right] + \left[\sum_{j=1}^{\infty}\psi^j R_{t+j}\right]. \tag{22}$$

Eq. (22) "is similar to the present discounted value formula for exchange rates derived in Engel and West (2005, 2006), and this equation states that any deviation of current exchange rates from their observed fundamentals should reflect the variation of the present discounted value of agent's expected future economic fundamentals (Balke et al., 2013, p. 3)."

3. Methodology

3.1. Emirmahmutoglu and Kose's (2011) test

It is well-known that the standard asymptotic theory is not applicable to hypothesis testing in level VARs if the variables are integrated or cointegrated (Sims et al., 1990; Toda and Phillips, 1993). To overcome this problem, Toda and Yamamoto (1995) have proposed an alternative approach for testing coefficient restrictions of a level VAR model for an integrated or cointegrated process. They recommend using a modified

Table 1
Summary of various monetary models of the exchange rate.

	$s_t = \alpha + \beta(m_t - m_t^*) + \phi(y_t - y_t^*) + \gamma(i_t - i_t^*) + \eta(\pi_t - \pi_t^*) + \varepsilon_t$			
	Money supply differential	Real output differential	Nominal interest rate differential	Expected inflation differential
The flexible price monetary model				
Bilson (1978a, 1978b)	$\beta = 1$	$\phi < 0$	$\gamma > 0$	$\eta = 0$
Frenkel (1976)	$\beta = 1$	$\phi < 0$	$\gamma = 0$	$\eta > 0$
The sticky price monetary model				
Dornbusch (1976)	$\beta = 1$	$\phi < 0$	$\gamma < 0$	$\eta = 0$
Frankel (1979)	$\beta = 1$	$\phi < 0$	$\gamma < 0$	$\eta > 0$

Source: Based on Frankel (1979). Term s_t is the log of the spot exchange rate, defined as the price of foreign currency in terms of domestic. Terms m_t and m_t^* are the logs of the domestic and foreign money supplies, respectively; y_t and y_t^* are the logs of domestic and foreign real income; and i_t and i_t^* the domestic and foreign interest rate. Terms π_t and π_t^* are the expected inflation rates, at home and abroad, respectively.

Wald (MWALD) test in a lag augmented VAR (LA-VAR) which has a conventional asymptotic chi-square distribution when a VAR($p + dmax$) is estimated, where p is the lag order and $dmax$ is the maximal order of integration suspected to occur in the process.

Emirmahmutoglu and Kose (2011) extend the LA-VAR approach via Meta analysis to test for Granger non-causality between variables in heterogeneous mixed panels. They consider the level VAR model with $p + dmax_i$ lags in heterogeneous mixed panels:

$$s_{i,t} = c_i + \sum_{m=1}^{p_{s_i}+dmax_i} \theta_{i,m} s_{i,t-m} + \sum_{m=1}^{p_{f_i}+dmax_i} \varphi_{i,m} f_{i,t-m} + \varepsilon_{i,t}, \tag{23}$$

$$f_{i,t} = \tilde{c}_i + \sum_{m=1}^{\tilde{p}_{f_i}+dmax_i} \tilde{\theta}_{i,m} f_{i,t-m} + \sum_{m=1}^{\tilde{p}_{s_i}+dmax_i} \tilde{\varphi}_{i,m} s_{i,t-m} + \epsilon_{i,t}, \tag{24}$$

where $i(=1,\ldots,N)$ denotes the i-th nation and $t(=1,\ldots,T)$ denotes the time index. The variables s_t and f_t, respectively, denote the logarithm of nominal exchange rates and market fundamentals. $p_{\cdot i}(=\{p_{s_i}, p_{f_i}\})$ and $\tilde{p}_{\cdot i}(=\{\tilde{p}_{s_i}, \tilde{p}_{f_i}\})$ are the lag lengths. The term $dmax_i$ is the maximal order of integration suspected to occur in the system for each i-th nation. For simplicity, we focus on testing Granger non-causality from f_t to s_t in Eq. (23) and from s_t to f_t in Eq. (24).

To test for Granger non-causality in this system, alternative causal relations are likely to be found for nation j: (i) There is one-way Granger causality from f_t (market fundamental) to s_t (nominal exchange rates) if the $\varphi_{j,1} = \varphi_{j,2} = \cdots = \varphi_{j,p_{f_j}} = 0$ does not hold, but the $\tilde{\varphi}_{j,1} = \tilde{\varphi}_{j,2} = \cdots = \tilde{\varphi}_{j,\tilde{p}_{s_j}} = 0$ holds.[10] (ii) There is one-way Granger causality from s_t (nominal exchange rates) to f_t (market fundamentals) if the $\varphi_{j,1} = \varphi_{j,2} = \cdots = \varphi_{j,p_{f_j}} = 0$ holds but $\tilde{\varphi}_{j,1} = \tilde{\varphi}_{j,2} = \cdots = \tilde{\varphi}_{j,\tilde{p}_{s_j}} = 0$ does not hold. This indicates that the present-value model of exchange rates is sustained through the uni-directional causality running from nominal exchange rates to market fundamentals. (iii) There is two-way Granger causality between nominal exchange rates and market fundamentals if neither the $\varphi_{j,1} = \varphi_{j,2} = \cdots = \varphi_{j,p_{f_j}} = 0$ nor the $\tilde{\varphi}_{j,1} = \tilde{\varphi}_{j,2} = \cdots = \tilde{\varphi}_{j,p_{s_j}} = 0$ holds. This result implies that the feedback hypothesis is linked. (iv) There is Granger non-causality between s_t and f_t if both the $\varphi_{j,1} = \varphi_{j,2} = \cdots = \varphi_{j,p_{f_j}} = 0$ and $\tilde{\varphi}_{j,1} = \tilde{\varphi}_{j,2} = \cdots = \tilde{\varphi}_{j,p_{s_j}} = 0$ hold. This is indicative of the *neutrality* hypothesis being sustained.

Emirmahmutoglu and Kose (2011) use the Fisher test statistic proposed by Fisher (1932) in order to test the Granger non-causality hypothesis in heterogeneous panels. Fisher (1932) considered combining several significant levels (p-values) of identical but independent tests. If the test statistics are continuous, the p-values p_i ($i = 1,\ldots,N$) are independent uniform (0,1) variables. In this case, the Fisher test statistic (λ) is written as follows:

$$\lambda = -2\sum_{i=1}^{N} \ln(p_i), \tag{25}$$

where p_i is the p-value corresponding to the Wald statistic of the i-th individual cross-section. This test statistic has a χ^2 distribution with 2N degrees of freedom. The test is valid only if N is fixed as $T \to \infty$.

However, the limit distribution of the Fisher test statistic is no longer valid in the presence of cross correlations among the cross-sectional units. As a way to deal with such inferential difficulty in panels with cross correlations, Emirmahmutoglu and Kose (2011) use the bootstrap methodology in their Granger non-causality test for cross-sectional dependent panels. In order to accommodate the contemporaneous correlation in panels, we obtain the empirical distribution of the test statistic by using the bootstrap method.[11] We consider the level VAR model with $p + dmax_i$ lags in heterogeneous mixed panels as shown in Eqs. (23) and (24). For readers' information, we briefly list the steps for conducting the bootstrapping panel Granger non-causality test for Eq (23) as follows.

Step 1: We determine the maximal order of integration of variables $dmax_i$ for each country by using the traditional unit root test such as the augmented Dickey-Fuller test. We then estimate Eq (23) by ordinary least square (OLS) and select the optimal lengths p_{s_i} and p_{f_i} via Schwarz information criteria (SBC) with maximum optimal length of six.
Step 2: Given the lag lengths p_{s_i}, p_{f_i}, and $dmax_i$ obtained from Step 1, we re-estimate Eq. (23) by OLS under the non-causality hypothesis, i.e., $\varphi_{i,1} = \varphi_{i,2} = \cdots = \varphi_{i,p_{f_i}} = 0$ and obtain their residuals:

[10] Note that the null hypothesis can be tested using a standard Wald statistic.

[11] Emirmahmutoglu and Kose (2011) provide an underpinning on the validity of the bootstrap used in this study. They propose a new panel causality approach based on Meta analysis in heterogeneous mixed panels because the Meta analysis approach has been efficiently applied to non-stationary heterogeneous panels by, for example, Maddala and Wu (1999) and Choi (2001). Emirmahmutoglu and Kose (2011) have investigated the finite sample properties of the causality test based on Meta analysis via Monte Carlo experiments in heterogeneous mixed panels. Their Monte Carlo results show that, under the cross-section dependency assumption, in general, the LA-VAR approach suffers from serious size distortion for small T. This problem is particularly serious if T is small as $N \to \infty$. On the contrary, the empirical size is converging at the 5% nominal size when fixed N as $T \to \infty$. The finite sample power of the LA-VAR approach, under the cross-section dependency, show that, in general, the LA-VAR approach performs satisfactory for whole values of T and N. Readers are also referred to Palm et al. (2011), Smeekes and Urbain (2014) and Westerlund and Smeekes (2018) for details on the bootstrap in panel data models with the cross-sectional dependence and I(1) variables.

$$\hat{\varepsilon}_{i,t} = s_{i,t} - \hat{c}_i - \sum_{m=1}^{p_{s_i}+dmax_i} \hat{\theta}_{i,m} s_{i,t-m} - \sum_{m=p_{f_i}+1}^{p_{f_i}+dmax_i} \hat{\varphi}_{i,m} f_{i,t-m}.$$

Step 3: We calculate the centered version of residuals suggested by Stine (1987):

$$\ddot{\varepsilon}_t = \hat{\varepsilon}_t - (T - p - d - 2)^{-1} \sum_{t=p+d+2}^{T} \hat{\varepsilon}_t$$

where $\hat{\varepsilon}_t = [\hat{\varepsilon}_{1,t}\ \hat{\varepsilon}_{2,t}\ \cdots\ \hat{\varepsilon}_{N,t}]'$, $p = \max(p_{f_i})$, and $d = \max(dmax_i)$. In order to preserve the cross covariance structure of the errors when bootstrapping, we stack the residuals for all countries such that $[\ddot{\varepsilon}_{i,t}]_{NT\times 1}$ and select randomly a full column with replacement at a time.

Step 4: We generate the bootstrap sample of $s_{i,t}$, i.e., $s^*_{i,t}$ from the bootstrapped residuals $\ddot{\varepsilon}^*_{i,t}$ under the null hypothesis:

$$s^*_{i,t} = \hat{c}_i + \sum_{m=1}^{p_{s_i}+dmax_i} \hat{\theta}_{i,m} s_{i,t-m} + \sum_{m=p_{f_i}+1}^{p_{f_i}+dmax_i} \hat{\varphi}_{i,m} f_{i,t-m} + \ddot{\varepsilon}^*_{i,t}$$

where the estimated $\hat{c}_i$, $\hat{\theta}_i$s, and $\hat{\varphi}_i$s are the estimations from Step 2.

Step 5: We substitute $s^*_{i,t}$ for $s_{i,t}$ in Eq. (23) and estimate it for each country and conduct Wald test to test for non-causality null hypothesis. The calculated Wald statistics have an asymptotic chi-square distribution with p_{f_i} degrees of freedom, and then we can compute individual p-values and thus the Fisher test statistic given in Eq. (25) is obtained.

3.2. Tests for cross-sectional dependence

One important issue to be considered in a panel data analysis is testing for cross-sectional dependency across series. For this purpose, we adopt three well-known statistics in the literature to test for cross-sectional dependence. First, we utilize the following Lagrange multiplier statistic for cross-sectional dependence (hereafter, CD_{BP}) developed by Breusch and Pagan (1980).

$$CD_{BP} = T \sum_{i=1}^{N-1} \sum_{j=i+1}^{N} \hat{\rho}_{ij}^2 \tag{26}$$

where $\hat{\rho}_{ij}$ is the estimated correlation coefficient among the residuals obtained from individual OLS estimations. Under the null hypothesis of no cross-sectional dependence with a fixed N and $T \to \infty$, CD_{BP} is asymptotically distributed as chi-squared with $N(N-1)/2$ degrees of freedom.

Pesaran (2004) indicates that the CD_{BP} test has a drawback when N is large, implying that it is not applicable when $N \to \infty$. To overcome this problem, the following Lagrange multiplier statistic for the cross-sectional dependence (hereafter, CD_{lm}) developed by Pesaran (2004) can be used.

$$CD_{lm} = \sqrt{\frac{1}{N(N-1)}} \sum_{i=1}^{N-1} \sum_{j=i+1}^{N} (T\hat{\rho}_{ij}^2 - 1) \tag{27}$$

Under the null hypothesis of no cross-sectional dependence with first $T \to \infty$ and then $N \to \infty$, this test statistic is asymptotically distributed as standard normal. However, this test is likely to exhibit substantial size distortions when N is large relative to T. A new test for the cross-sectional dependence (hereafter, CD_P) of Pesaran (2004) can be used where N is large and T is small. The CD_P statistic is calculated as follows:

$$CD_P = \sqrt{\frac{2T}{N(N-1)}} \left(\sum_{i=1}^{N-1} \sum_{j=i+1}^{N} \hat{\rho}_{ij} \right) \tag{28}$$

Under the null hypothesis of no cross-sectional dependence with $T \to \infty$ and $N \to \infty$ in any order, the CD_P test is asymptotically distributed as standard normal.

3.3. Tests for slope homogeneity

Determining whether slope coefficients are homogeneous or heterogeneous is also important in a panel non-causality analysis to impose causality restrictions on estimated coefficients. As noted by Granger (2003), imposing the joint restriction for the whole panel is a very strong null hypothesis. Moreover, Breitung (2005) also pointed out that the homogeneity assumption for the parameters is unable to capture heterogeneity due to nation-specific characteristics. By letting the parameter vector $\boldsymbol{b} = (\beta, \phi, \gamma, \eta)'$ in Eq. (13), the null hypothesis of slope homogeneity ($H_0 : \boldsymbol{b}_i = \boldsymbol{b}$, for all i) is tested against the alternative hypothesis of heterogeneity ($H_1 : \boldsymbol{b}_i \neq \boldsymbol{b}$) for a non-zero fraction of pair-wise slopes for $i \neq j$.

The standard Wald test is widely used in testing for the null hypothesis of slope homogeneity. However, as pointed out by Pesaran et al. (2008), the test based on the Wald principle is applicable only for cases in which the cross-sectional dimension (N) is relatively small, the time dimension (T) of the panel is large, the explanatory variables are strictly exogenous, and the error variances are homoskedastic. To overcome these problems, Pesaran et al. (2008) developed a standardized version of Swamy's test (Swamy, 1970) for testing slope homogeneity in large panels. Swamy's test is valid when $(N, T) \to \infty$ without any restrictions on the relative expansion rates of N and T when the error terms are normally distributed. The Swamy test for slope homogeneity is:

$$\tilde{S} = \sum_{i=1}^{N} \left(\hat{\boldsymbol{b}}_i - \hat{\boldsymbol{b}}_{WFE}\right)' \frac{\boldsymbol{x}_i' \boldsymbol{M}_\tau \boldsymbol{x}_i}{\hat{\sigma}_i^2} \left(\hat{\boldsymbol{b}}_i - \hat{\boldsymbol{b}}_{WFE}\right), \tag{29}$$

where $\hat{\boldsymbol{b}}_i$ is the pooled OLS estimator; $\hat{\boldsymbol{b}}_{WFE}$ is the weighted fixed effect pooled estimator; $\boldsymbol{M}_\tau = \boldsymbol{I}_T - \boldsymbol{Z}_i(\boldsymbol{Z}_i'\boldsymbol{Z}_i)^{-1}\boldsymbol{Z}_i'$ and $\boldsymbol{Z}_i = (\tau_T, \boldsymbol{x}_i)$, where τ_T is a $T \times 1$ vector of ones; $\boldsymbol{x}_i = (s_t, f_t)$, and $f_t = ((m_t - m_t^*), (y_t - y_t^*), (i_t - i_t^*), (\pi_t - \pi_t^*))$; and $\hat{\sigma}_i^2$ is the estimator of the error variance σ_i^2.

In the case where N is fixed and $T \to \infty$, the $\tilde{S}$ test has an asymptotic chi-squared distribution with $k(N-1)$ degrees of freedom where k is the number of explanatory variables.[12] The standardized dispersion statistic is as follows:

$$\tilde{\Delta} = \sqrt{N} \left(\frac{N^{-1}\tilde{S} - k}{\sqrt{2k}} \right). \tag{30}$$

Under the null hypothesis with the condition of $(N, T) \to \infty$, as long as $\sqrt{N}/T \to \infty$ and the error terms are normally distributed, the $\tilde{\Delta}$ test has an asymptotic standard normal distribution. The $\tilde{\Delta}$ test can be improved for small samples by using the following bias-adjusted version:

$$\tilde{\Delta}_{adj} = \sqrt{N} \left(\frac{N^{-1}\tilde{S} - k}{\sqrt{2k(T-k-1)/T+1}} \right). \tag{31}$$

4. Data and results

4.1. Data descriptions

This study aims to assess if the findings of Engel and West (2005), Bahmani-Oskooee et al. (2015) and Ko and Ogaki (2015) are applicable or diverse using the same data sets, sample periods and currency, but alternative approach. Hence, we use the same dataset as in Engel and West (2005), but the difference is that we suggest using the panel data with bootstrap developed by Emirmahmutoglu and Kose (2011) as

[12] In order to save space, we refer to Pesaran et al. (2008) for the details of Swamy's test and the estimators described in Eq. (29).

Table 2
Results of the ADF unit root test.

	s			$(m-m^*)$			$(y-y^*)$		
	Level	1st Difference	2nd Difference	Level	1st Difference	2nd Difference	Level	1st Difference	2nd Difference
Canada	0.764	0.050**	–	1.000	0.000***	–	0.123	0.000***	–
France	0.749	0.000***	–	0.409	0.018**	–	0.248	0.000***	–
Germany	0.717	0.000***	–	0.998	0.000***	–	0.930	0.000***	–
Italy	0.595	0.000***	–	0.100	0.000***	–	0.532	0.005***	–
Japan	0.634	0.000***	–	0.993	0.000***	–	0.998	0.000***	–
UK	0.699	0.000***	–	0.207	0.002***	–	0.114	0.000***	–
	$(i-i^*)$			$(\pi-\pi^*)$			$(m-m^*)-(y-y^*)$		
	Level	1st Difference	2nd Difference	Level	1st Difference	2nd Difference	Level	1st Difference	2nd Difference
Canada	0.455	0.000***	–	0.959	0.000***	–	1.000	0.000***	–
France	0.002***	–	–	0.946	0.001***	–	0.671	0.007***	–
Germany	0.919	0.000***	–	0.063	0.000***	–	0.999	0.000***	–
Italy	0.000***	–	–	0.616	0.479	0.005***	0.277	0.000***	–
Japan	0.085*	0.000***	–	0.005***	–	–	1.000	0.000***	–
UK	0.064*	0.000***	–	0.535	0.013**	–	0.112	0.000***	–

The symbols ***, **, and * denote rejection at the significance levels of 1%, 5%, and 10%, respectively.

opposed to the standard Johansen approach. The sample exchange rates include the US dollar against the Canadian dollar, the French franc, the Deutsche mark, the Italian lira, the Japanese yen, and the British pound, covering quarterly data from 1974Q1 to 2001Q3. Therefore, the data contain six countries ($N = 6$) with sample size 111 observations ($T = 111$). Based upon the flexible price monetary model (cf. Eq. (4)) and the sticky price monetary model (cf. Eq. (13)) reviewed in Section 2, we consider the following five macroeconomic fundamentals to implement the empirical study. They are the money supply fundamental $(m_t - m_t^*)$, the output fundamental $(y_t - y_t^*)$, the PPP fundamental $(\pi_t - \pi_t^*)$, the UIP fundamental $(i_t - i_t^*)$ as well as the monetary fundamental $(m_t - m_t^*) - \phi(y_t - y_t^*)$. Following Mark (1995) and Rapach and Wohar (2002), we use a fixed value of $\phi = 1$ for the income elasticity in the money demand so the monetary fundamental is $(m_t - m_t^*) - (y_t - y_t^*)$. For more details on the data see Engel and West (2005).

The steps for conducting Emirmahmutoglu and Kose's (2011) panel Granger non-causality analysis is as follows. First of all, we start by testing for the integrated properties of the series by using the individual and panel unit root tests. Next, as highlighted by Bai and Kao (2006), testing for the cross-sectional dependence in a panel non-causality study is crucial for selecting the appropriate estimator. Therefore, in the second step, we test for the cross-sectional dependence of the data and hypothesis of slope homogeneity. Finally, we estimate the LA-VAR model (cf. Eqs. (23) and (24)) and conduct the panel Granger non-causality tests with bootstrap critical values from the market fundamentals to nominal exchange rates in Eq. (23) and from nominal exchange rates to the market fundamentals in Eq. (24).

4.2. Results of the individual and panel unit root tests

The first step in conducting the Emirmahmutoglu and Kose (2011) test is to investigate the integrated properties of the series for all countries. Hence, Table 2 report the augmented Dickey-Fuller (ADF) unit root test on the level, first difference and second difference of the series. As a consequence of the ADF unit root test, with the exceptions of France, the maximum order of integration in the LA-VAR system is determined as one for these countries. For the purpose of robustness, we employ the panel unit root tests of Im et al. (2003) (IPS hereafter) and Maddala and Wu (1999) for detecting the degree of integration of empirical variables in this paper. These tests enable us to recognize that there may be a mixture of stationary and nonstationary processes in the panel under the alternative hypothesis. We present the test results in Table 3. With the exception of $(\pi_t - \pi_t^*)$, it is found that the null hypothesis of a unit root cannot be rejected for the level data of all empirical series based on the IPS and the Maddala and Wu (1999) PP-Fisher χ^2 test at the 5 percent significance level.[13] Moreover, for the first-difference of the data, the null hypothesis of a unit root must be rejected at the 5 percent significance level for all series.

In order to consider the cross-sectional dependence (CSD) in testing for the panel unit root,[14] we employ the heteroskedasticity-robust panel unit root tests proposed by Herwartz and Siedenburg (2008), Demetrescu and Hanck (2012) and Herwartz et al. (2017). While the former two tests are robust to time-varying volatility when the data contain only an intercept, the latter test is unique because it is asymptotically pivotal for trending heteroskedastic panels. As noted by Herwartz et al. (2018, p. 185), "the test in Herwartz et al. (2017) is unique in avoiding complicated resampling techniques used to define panel unit root tests in trending heteroskedastic panels. This is an important property because several macroeconomic time series, including gross domestic product, money supply, and commodity prices, contain a linear trend (Westerlund, 2015)." We summarize the heteroskedasticity-robust panel unit root test results in Table 4.[15] It is found that the null hypothesis of a unit root cannot be rejected for the level data of all empirical series based on the statistics of t_{HS} (Herwartz and Siedenburg, 2008), t_{DH} (Demetrescu and Hanck, 2012) and t_{HMW} (Herwartz et al., 2017) at the 5 percent significance level. For the first-difference of the data, with the exceptions of $(i_t - i_t^*)$ and $(\pi_t - \pi_t^*)$, the results of the t_{HS}, t_{DH} and t_{HMW} show that the null hypothesis of a unit root must be rejected at the 5 percent significance level for s_t, $(m_t - m_t^*)$ and $(y_t - y_t^*)$.

The results of the panel unit roots are in line with Bahmani-Oskooee et al. (2015). By using the augmented Dickey-Fuller unit root test, they also find that not all empirical variables of level data are I(1) series. Therefore, they argue that the weak cointegration results reported in Engel and West (2005) could be result from using the Johansen (1991) approach with a combination of variables being I(1) and I(0). Hence, Bahmani-Oskooee et al. (2015) instead adopt the ARDL approach of Pesaran et al. (2001) to avoid the above problem.

[13] Im et al. (2003) point out that due to the heterogeneous nature of the alternative hypothesis in their test, caution has to be exercised when interpreting results because the null hypothesis of a unit root in each cross section may be rejected when only a fraction of the series in the panel is stationary. An additional concern here is that the presence of cross-sectional dependencies can undermine the asymptotic normality of the IPS test and lead to over-rejection of the null hypothesis of joint non-stationarity.

[14] We thank an anonymous referee for raising this issue to us.

[15] Readers are referred to Herwartz et al. (2018) for detailed descriptions of the t_{HS}, t_{DH} and t_{HMW} statistics and the command `xtpurt` in Stata, which implements the heteroskedasticity-robust panel unit root tests.

Table 3
Results of the panel unit root tests.

	Level		1st Difference	
	statistics	P-value	statistics	P-value
Panel (A): s_t				
IPS	−1.076	0.141	−10.878	0.000[c]
Maddala and Wu (1999) ADF-Fisher Chi Square	13.849	0.311	125.045	0.000[c]
Maddala and Wu (1999) PP-Fisher Chi Square	8.382	0.755	310.827	0.000[c]
Panel (B): $(m_t - m_t^*)$				
IPS	1.687	0.954	−6.035	0.000[c]
Maddala and Wu (1999) ADF-Fisher Chi Square	9.661	0.646	58.105	0.000[c]
Maddala and Wu (1999) PP-Fisher Chi Square	5.328	0.946	203.424	0.000[c]
Panel (C): $(y_t - y_t^*)$				
IPS	−0.0415	0.484	−24.542	0.000[c]
Maddala and Wu (1999) ADF-Fisher Chi Square	14.038	0.298	332.934	0.000[c]
Maddala and Wu (1999) PP-Fisher Chi Square	13.749	0.317	336.485	0.000[c]
Panel (D): $(i_t - i_t^*)$				
IPS	0.217	0.586	−5.978	0.000[c]
Maddala and Wu (1999) ADF-Fisher Chi Square	15.046	0.239	66.625	0.000[c]
Maddala and Wu (1999) PP-Fisher Chi Square	19.324	0.081[a]	249.013	0.000[c]
Panel (E): $(\pi_t - \pi_t^*)$				
IPS	−1.751	0.040[b]	−19.046	0.000[c]
Maddala and Wu (1999) ADF-Fisher Chi Square	21.360	0.045[b]	247.014	0.000[c]
Maddala and Wu (1999) PP-Fisher Chi Square	69.359	0.000[c]	356.465	0.000[c]

[a] Denotes rejection at the significance level of 10%.
[b] Denotes rejection at the significance level of 5%.
[c] Denotes rejection at the significance level of 1%.

As a consequence of the panel unit root test, the maximum order of integration in the LA-VAR system is determined as one or two of these nations. Remember that *dmax* is the maximal order of integration suspected to occur in the process. The only prior information needed for the LA-VAR approach is the maximum order of integration of the processes. We use $dmax = 1$ or 2 for these nations in the panel specification.

4.3. Results of the panel Granger causality

Following Engel and West (2005) and Bahmani-Oskooee et al. (2015), we estimate five groups of bivariate panel models between nominal exchange rates and every fundamentals. The lag order for the bivariate panel model is determined by the Akaike information criterion (AIC) and the results show that the optimal lag order is equal to

Table 4
Results of the panel unit root tests (cont.).

	Level		1st Difference		2nd Difference	
	statistics	*p*-value	statistics	*p*-value	statistics	*p*-value
Panel (A): s_t						
Herwartz and Siedenburg (2008) t_{HS}	0.695	0.757	−2.507	0.006[c]	–	–
Demetrescu and Hanck (2012) t_{DH}	2.021	0.978	−2.943	0.002[c]	–	–
Herwartz et al. (2017) t_{HMW}	0.982	0.837	−1.615	0.053[a]	–	–
Panel (B): $(m_t - m_t^*)$						
Herwartz and Siedenburg (2008) t_{HS}	1.038	0.850	−2.649	0.004[c]	–	–
Demetrescu and Hanck (2012) t_{DH}	−0.884	0.188	−2.993	0.001[c]	–	–
Herwartz et al. (2017) t_{HMW}	2.102	0.982	−2.471	0.007[c]	–	–
Panel (C): $(y_t - y_t^*)$						
Herwartz and Siedenburg (2008) t_{HS}	1.317	0.906	−2.474	0.007[c]	–	–
Demetrescu and Hanck (2012) t_{DH}	0.234	0.592	−2.485	0.007[c]	–	–
Herwartz et al. (2017) t_{HMW}	1.593	0.944	−2.167	0.015[b]	–	–
Panel (D): $(i_t - i_t^*)$						
Herwartz and Siedenburg (2008) t_{HS}	0.837	0.799	−0.781	0.217	−2.653	0.004[c]
Demetrescu and Hanck (2012) t_{DH}	1.107	0.866	−1.091	0.138	−2.571	0.005[c]
Herwartz et al. (2017) t_{HMW}	0.124	0.549	−1.192	0.117	−2.317	0.010[c]
Panel (E): $(\pi_t - \pi_t^*)$						
Herwartz and Siedenburg (2008) t_{HS}	1.135	0.872	−0.398	0.345	−3.024	0.001[c]
Demetrescu and Hanck (2012) t_{DH}	2.486	0.994	−0.330	0.371	−3.688	0.000[c]
Herwartz et al. (2017) t_{HMW}	0.720	0.764	−0.830	0.203	−2.523	0.006[c]

[a] Denotes rejection at the significance level of 10%.
[b] Denotes rejection at the significance level of 5%.
[c] Denotes rejection at the significance level of 1%.

four. Engel and West (2005) and Ko and Ogaki (2015) also estimate the bivariate VAR model with the lag order equals to four and conduct the Granger non-causality tests accordingly. For the purpose of robustness, in this study, we also consider to estimate the bivariate panel model with lag order equals to 1, …, 4, respectively. All of the empirical results of the bivariate panel models are summarized in Tables 5 and 6.[16]

One important issue to be considered in a panel data analysis is testing for the cross-sectional dependency across nations. Basher and Westerlund (2009) suggest that accounting for the effects of cross-section dependence is crucial in the analysis of the monetary model in a panel framework. Following Kar et al. (2011), we carry out three different tests, i.e., CD_{BP}, CD_{lm} and CD_P, to investigate the existence of the cross-sectional dependence. We summarize the results in the panel (B) of Tables 5 and 6. It is clear that the null hypothesis of no cross-sectional dependence across the members of the panel is strongly rejected at the one percent significance level for the bivariate panel model, indicating that the bootstrap critical value is required in conducting the panel Granger non-causality test. This finding implies that a shock occurring in one country seems to be transmitted to other countries. Hence, there seems to be evidence of cross-sectional dependence in the data.

Panel (C) of Tables 5 and 6 reports the results of the slope homogeneity test for the bivariable panel model. Both tests ($\widetilde{\Delta}$ and $\widetilde{\Delta}_{adj}$) reject the null hypothesis of the slope homogeneity at the 1% significance level, supporting the view that the parameters are heterogeneous. This finding simply implies that the panel non-causality analysis by imposing the homogeneity restriction on the variable of interest results in misleading inferences. In this respect, the panel non-causality analysis based on estimating a panel vector autoregression and/or panel vector error correction model by means of the generalized method of moments and the pooled ordinary least squares estimator is not an appropriate approach in detecting causal linkages between exchange rates and macroeconomic fundamentals. In order to account for this we use the bootstrapped *p*-values in conducting the panel Granger non-causality test.

The final step is to perform the LA-VAR approach using mixed panels to test the hypothesis that there is a relationship between nominal exchange rates and macroeconomic fundamentals using the level data. Given that exchange rate fundamentals are highly persistent and the discount factor is very close to a one, Engel and West (2005) argue that exchange rates are the present discounted value of expected economic fundamentals and their stochastic behavior look like a random walk. An important empirical implication is that current exchange rates are helpful for predicting the future fundamentals in terms of Granger's definition. Following Engel and West (2005), we perform the Granger non-causality tests based on the bivariate panel model. We use five macroeconomic fundamentals and therefore there are five groupings: $(s_t, (m_t - m_t^*))$, $(s_t, (y_t - y_t^*))$, $(s_t, (\pi_t - \pi_t^*))$, $(s_t, (i_t - i_t^*))$ and $(s_t, (m_t - m_t^*) - (y_t - y_t^*))$. An advantage of the LV-VAR procedure is that it is applied irrespective of whether the variables are I(1) or I(0) process. Hence, all variables are entered in level data rather than in first-difference data. We consider the lag order from 1 to 4 and, therefore, there are totally 20 groups. The results of the LA-VAR approach are given in panel (A) of Tables 5 and 6. In each case, we perform the Granger non-causality tests for the null hypothesis that nominal exchange rates does not Granger-cause each of the fundamentals ($H_0 : s_t \nRightarrow f_t$), and the null hypothesis that each of the fundamentals does not Granger-cause nominal exchange rates ($H_0 : f_t \nRightarrow s_t$).

From Table 5, the null hypothesis of nominal exchange rate does not Granger-cause the money supply fundamental ($H_0 : s_t \nRightarrow (m_t - m_t^*)$) is rejected at the conventional significance level for every country except Germany. This is also true for the case of the UIP fundamental for France, Germany, Italy, Japan and the UK because the null hypothesis of Granger non-causality running from exchange rates to the interest differentials ($H_0 : s_t \nRightarrow (i_t - i_t^*)$) is strongly rejected. With the exceptions of Germany and the UK, the null hypothesis of nominal exchange rate does not Granger-cause the monetary fundamental, i.e., $H_0 : s_t \nRightarrow (m_t - m_t^*) - (y_t - y_t^*)$, is rejected at the five percent significance level for other countries. In general, the empirical evidences show that nominal exchange rates are helpful for predicting the future behaviors of the money supply, the interest differential and the monetary fundamentals, which is in support of the present-value model of exchange rates proposed by Engel and West (2005). However, for the output fundamental, $(y_t - y_t^*)$, the unidirectional Granger non-causality running from nominal exchange rate to the output fundamental is not rejected at the five percent significance level for every country. With the exception of Italy, the empirical evidences show that the nominal exchange rate is also not helpful in forecasting the PPP fundamental since the null hypothesis of $H_0 : s_t \nRightarrow (\pi_t - \pi_t^*)$ is not rejected at the 5% significance level.

Panel (A) of Table 5 also shows the Fisher test statistic values combined with the *p*-values to assess an overall hypothesis for the six countries.[17] This test statistic λ is distributed as χ^2_{2N} under the cross-section independency assumption. As shown in Table 5, the null hypothesis of the cross-section independence assumption is strongly rejected and, therefore, the limit distribution of the Fisher test statistic is no longer valid. In the presence of the cross-section dependence in mixed panels, we apply the bootstrap method to generate the empirical distributions of the Fisher test. The bootstrap distribution of the Fisher test statistics is derived from 10,000 replications. Bootstrap critical values are obtained at the 10%, 5% and 1% levels based on these empirical distributions. The empirical results show that, to our surprise, the λ statistics are not significantly at the conventional level for the most of the macroeconomic fundamentals. It is significance at the five percent level only for the null hypothesis of $H_0 : s_t \nRightarrow (\pi_t - \pi_t^*)$.

Panel (A) of Table 6 summarizes the results of Granger non-causality test running from each of the fundamentals to nominal exchange rates under the bivariate panel model. The optimal lag order for the bivariate panel model is equal to four based on the AIC information criterion. Remember that we also consider the other model specifications with lag order equals to 1, …,4 for the purpose of robustness. We first carry out three different tests, i.e., CD_{BP}, CD_{lm} and CD_P, to investigate the existence of the cross-sectional dependence. As shown in panel (B) of Table 6, the test results show that the null hypothesis of no cross-sectional dependence across the members of the panel are strongly rejected at the one percent significance level for the bivariate panel model, indicating that the bootstrap critical value is required in conducting the panel Granger non-causality test. Next, we examine the slope homogeneity test for the bivariable panel model and summarize the results in the panel (C) of Table 6. Both tests ($\widetilde{\Delta}$ and $\widetilde{\Delta}_{adj}$) reject the null hypothesis of the slope homogeneity at the 1% significance level, supporting the view that the parameters are heterogeneous.

Finally, we turn our attention to test for the null hypothesis of Granger non-causality running from every fundamentals to exchange rates ($H_0 : f_t \nRightarrow s_t$). The empirical results of the Granger non-causality tests show that, with the exception of Canada, there is a unidirectional Granger causality running from the money supply fundamental, $(m_t - m_t^*)$, to nominal exchange rates for each country. This is also true for the case of the output fundamental $(y_t - y_t^*)$. For the cases of the PPP fundamental $(\pi_t - \pi_t^*)$, the UIP fundamental $(i_t - i_t^*)$ and the monetary fundamental $(m_t - m_t^*) - (y_t - y_t^*)$, they are helpful for predicting the future behavior of nominal exchange rates for each country based

[16] The empirical results of the Granger non-causality tests for the bivariate panel model with lag order equals to 1, 2, and 3 are available from the authors upon request.

[17] We code the panel Granger non-causality by using the Winrats software according to the Matlab code provided by Professor Emirmahmutoglu. We thank him for providing us with his Matlab code for reference.

Table 5

Granger causality tests for H_o: exchange rates do not Granger cause fundamentals – bivariate model with lag order equals to 4.

	$H_o: s \nRightarrow (m-m^*)$ (1978Q1 ~1998Q4)			$H_o: s \nRightarrow (y-y^*)$ (1974Q1 ~2001Q3)			$H_o: s \nRightarrow (i-i^*)$ (1978Q3 ~2001Q3)			$H_o: s \nRightarrow (\pi-\pi^*)$ (1974Q1 ~2001Q3)			$H_o: s \nRightarrow (m-m^*)-(y-y^*)$ (1978Q1 ~1998Q4)		
	Wald Statistics	Lags	*dmax*	Wald Statistics	Lags	*dmax*	Wald Statistics	Lags	*dmax*	Wald Statistics	Lags	*dmax*	Wald Statistics	Lags	*dmax*
Emirmahmutoglu and Kose's (2011) Granger causality test															
Canada	33.122***	4	1	3.907	4	1	5.800	4	1	0.031	4	1	23.985***	4	1
	(0.000)			(0.419)			(0.215)			(1.000)			(0.000)		
France	41.550***	4	1	1.456	4	1	62.622***	4	1	1.284	4	2	48.335***	4	1
	(0.000)			(0.834)			(0.000)			(0.864)			(0.000)		
Germany	5.036	4	1	1.507	4	1	9.931**	4	1	4.143	4	1	6.053	4	1
	(0.284)			(0.825)			(0.042)			(0.387)			(0.195)		
Italy	30.633***	4	1	1.592	4	1	21.780***	4	1	10.084**	4	1	35.413***	4	1
	(0.000)			(0.810)			(0.000)			(0.039)			(0.000)		
Japan	25.207***	4	1	3.140	4	1	25.631***	4	1	5.623	4	1	39.033***	4	1
	(0.000)			(0.535)			(0.000)			(0.229)			(0.000)		
UK	11.700**	4	1	3.844	4	1	13.140**	4	1	2.320	4	1	5.165	4	1
	(0.020)			(0.428)			(0.011)			(0.677)			(0.271)		
λ statistics	118.186			5.859			111.408			12.405*			129.165		
Bootstrap critical values of λ test															
10%	128.557			31.102			154.983			9.139			203.091		
5%	162.229			41.751			204.117			12.448			256.955		
1%	238.697			70.476			332.561			21.075			380.400		

	$(m-m^*)$	$(y-y^*)$	$(i-i^*)$	$(\pi-\pi^*)$	$(m-m^*)-(y-y^*)$
(B) Results of the cross-sectional dependence					
CD_{BP}	119.506***	387.959***	420.022***	148.459***	111.062***
CD_{lm}	19.080***	68.093***	73.947***	24.366***	17.538 ***
CD_p	9.650***	18.852***	19.700***	11.111***	11.158 ***
(C) Testing results of slope homogeneity					
$\widetilde{\Delta}$	12.526***	20.532***	7.050***	36.011***	16.456***
$\widetilde{\Delta}_{adj}$	12.753***	20.812***	7.165***	36.504***	16.754***

Note for Panel (A): Number in parentheses is p-value. Lag orders k_i are selected by minimizing the Akaike information criteria. *Note* for Panel (B): The critical value of CD_{BP} at the 1% level is 30.578. The critical value of CD_{lm} and CD_P at the 1% level is 2.575. *Note* for Panel (C): The critical value of $\widetilde{\Delta}$ and $\widetilde{\Delta}_{adj}$ at the 1% level is 2.575. Symbols $*$, $**$, and $***$, respectively, denote rejection at the significance level of 10%, 5% and 1%.

Table 6
Granger causality tests for H_o: fundamentals do not Granger cause exchange rates – bivariate model with lag order equals to 4.

	$H_o: (m - m^*) \nRightarrow s$ (1978Q1 ~1998Q4)			$H_o: (y - y^*) \nRightarrow s$ (1974Q1 ~2001Q3)			$H_o: (i - i^*) \nRightarrow s$ (1978Q3 ~2001Q3)			$H_o: (\pi - \pi^*) \nRightarrow s$ (1974Q1 ~2001Q3)			$H_o: (m - m^*) - (y - y^*) \nRightarrow s$ (1978Q1 ~1998Q4)		
	WaldStatistics	Lags	*dmax*	WaldStatistics	Lags	*dmax*	WaldStatistics	Lags	*dmax*	WaldStatistics	Lags	*dmax*	WaldStatistics	Lags	*dmax*
(A) Emirmahmutoglu and Kose's (2011) Granger causality test															
Canada	7.340	4	1	6.177	4	1	12.302**	4	1	42.951***	4	1	10.333**	4	1
	(0.119)			(0.186)			(0.015)			(0.000)			(0.035)		
France	135.608***	4	1	111.954***	4	1	137.669***	4	1	80.299***	4	2	195.024***	4	1
	(0.000)			(0.000)			(0.000)			(0.000)			(0.000)		
Germany	424.637***	4	1	60.027***	4	1	108.573***	4	1	81.919***	4	1	87.502***	4	1
	(0.000)			(0.000)			(0.000)			(0.000)			(0.000)		
Italy	143.531***	4	1	214.260***	4	1	99.580***	4	1	58.020***	4	1	205.790***	4	1
	(0.000)			(0.000)			(0.000)			(0.000)			(0.000)		
Japan	259.147***	4	1	27.425***	4	1	343.238***	4	1	80.527***	4	1	189.747***	4	1
	(0.000)			(0.000)			(0.000)			(0.000)			(0.000)		
UK	57.786***	4	1	115.505***	4	1	156.951***	4	1	18.025***	4	1	133.430***	4	1
	(0.000)			(0.000)			(0.000)			(0.001)			(0.000)		
λ statistics	980.665			494.685			810.95			321.758			774.561		
Bootstrap critical values of λ test															
10%	1216.086			1159.341			1178.702			1127.655			1217.342		
5%	1493.055			1421.998			1452.352			1397.534			1498.574		
1%	2119.606			2023.876			2053.869			1993.798			2118.962		

	$(m - m^*)$	$(y - y^*)$	$(i - i^*)$	$(\pi - \pi^*)$	$(m - m^*) - (y - y^*)$
(B) Results of the cross-sectional dependence					
CD_{BP}	307.798***	400.298***	354.18***	462.412***	295.445***
CD_{lm}	53.457***	70.345***	61.93***	81.686***	51.202 ***
CD_p	14.411***	18.315***	15.45***	16.909***	16.468 ***
(C) Testing results of slope homogeneity					
$\widetilde{\Delta}$	21.037***	16.772***	6.076***	21.062***	25.782***
$\widetilde{\Delta}_{adj}$	21.419***	17.001***	6.175***	21.349***	26.249***

Note for Panel (A): Number in parentheses is p-value. Lag orders k_i are selected by minimizing the Akaike information criteria. *Note* for Panel (B): The critical value of CD_{BP} at the 1% level is 30.578. The critical value of CD_{lm} and CD_P at the 1% level is 2.575. *Note* for Panel (C): The critical value of $\widetilde{\Delta}$ and $\widetilde{\Delta}_{adj}$ at the 1% level is 2.575. Symbols $*$, $**$, and $***$, respectively, denote rejection at the significance level of 10%, 5% and 1%.

Table 7
Granger causality tests for H_o: fundamentals do not Granger cause exchange rates – multivariate model with lag order equals to 4.

	H_o : $(m-m^*) \nRightarrow s$	H_o : $(y-y^*) \nRightarrow s$	H_o : $(i-i^*) \nRightarrow s$	H_o : $(\pi-\pi^*) \nRightarrow s$	H_o : $(m-m^*) \nRightarrow s$ $(y-y^*) \nRightarrow s$	H_o : $(m-m^*) \nRightarrow s(y-y^*) \nRightarrow s$ $(i-i^*) \nRightarrow s(\pi-\pi^*) \nRightarrow s$	Lags	*dmax*
(A) Emirmahmutoglu and Kose's (2011) Granger causality test, (1978Q3 ~1998Q4)								
Canada	7.153	20.167***	18.415***	38.526***	27.438***	75.209***	4	1
	(0.128)	(0.000)	(0.001)	(0.000)	(0.001)	(0.000)		
France	343.936***	458.283***	516.998***	220.874***	698.625***	992.687***	4	1
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)		
Germany	294.594***	110.842***	61.166***	45.759***	394.952***	527.587***	4	1
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)		
Italy	174.462***	173.592***	231.021***	127.673***	363.930***	706.684***	4	2
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)		
Japan	185.795***	150.548***	282.068***	348.740***	351.936***	1015.979***	4	1
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)		
UK	65.758***	76.291***	75.296***	40.142***	141.758***	353.824***	4	1
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)		
λ statistics	1023.261**	941.016*	1135.558**	775.150	1829.116**	3316.079*		
Bootstrap critical values of λ test								
10.00%	810.699	795.225	822.143	784.781	1487.589	3259.204		
5.00%	1004.357	986.055	1016.961	975.087	1753.983	3676.292		
1.00%	1445.893	1396.573	1432.748	1374.370	2336.145	4576.212		
(B) Results of the cross-sectional dependence								
CD_{BP}	218.306***							
CD_{lm}	37.118***							
CD_p	10.652***							
(C) Testing results of slope homogeneity								
$\widetilde{\Delta}$	16.796***							
$\widetilde{\Delta}_{adj}$	17.439***							

Note for Panel (A): Number in parentheses is *p*-value. Lag orders k_i are selected by minimizing the Akaike information criteria. *Note* for Panel (B): The critical value of CD_{BP} at the 1% level is 30.578. The critical value of CD_{lm} and CD_P at the 1% level is 2.575. *Note* for Panel (C): The critical value of $\widetilde{\Delta}$ and $\widetilde{\Delta}_{adj}$ at the 1% level is 2.575. Symbols $*$, $**$, and $***$, respectively, denote rejection at the significance level of 10%, 5% and 1%.

on the fact that the null hypothesis of Granger non-causality running from the fundamentals to exchange rates ($H_0 : f_t \nRightarrow s_t$) is rejected at the conventional significance level. However, again, the λ statistics are not significant at the conventional level for all of the fundamentals, which is contradicted by the results of individual Granger non-causality tests.

The inconsistency between the results of individual Granger non-causality test and the λ statistic between exchange rates and fundamentals, to the best of authors' knowledge, might result from the problem of 'omitted variable' by estimating the bivariate panel model. In order to validate this possibility, we estimate a multivariate panel model for nominal exchange rates and the whole set of economic fundamentals. That is, we consider to estimate Eqs. (23) and (24) with the variables of $[s_t, (m_t - m_t^*), (y_t - y_t^*), (i_t - i_t^*), (\pi_t - \pi_t^*)]$. Again, for the purpose of robustness, we consider to estimate the multivariate panel model with the lag order equals to 1, …, 4, respectively. The optimal lag order for the multivariate panel model is equal to four and we summarize the results in Table 7.[18] Panels (B) and (C) of Table 7 show that the hypotheses of the cross-sectional dependence and the slope homogeneity are rejected at the 1% significant level. From panel (A) of Table 7, it shows that all of the fundamentals are helpful for predicting nominal exchange rates because the null hypothesis of Granger non-causality running from the fundamentals to exchange rates is rejected at one percent significance level. Besides, with the exception of the PPP fundamental $(\pi_t - \pi_t^*)$, the λ statistics are significantly at the conventional level for all of the macroeconomic fundamentals, implying that the null hypothesis of Granger non-causality running from the fundamentals to exchange rates is rejected. As such, the results of the λ statistics are in line with the results of individual Granger non-causality tests.

4.4. Comparison of results with selected literature

Bahmani-Oskooee et al. (2015) and Ko and Ogaki (2015) have devoted their efforts to the issue of the exchange rate disconnect puzzle using the same data in Engel and West (2005). They adopt the time series approach of cointegration and the Granger non-causality to dissect the causal relations between nominal exchange rates and market fundamentals. However, a common weakness of these studies is that they do not take account of the cross-sectional dependence of the data. This paper fills this gap by applying the panel model with bootstrap techniques to overcome the low power problem of traditional approaches.

For readers' information, we summarize their empirical findings and ours in Table 8 as follows. By using the Johansen (1991) cointegration approach, Engel and West (2005) find almost no evidence of cointegration in the multivariate case, and only 5 out of 24 cases in the bivariate cases and find that there are significant Granger causality relations from exchange rates to fundamentals, and they find no evidence of causality running from the market fundamentals to exchange rates. Ko and Ogaki (2015) claim that the findings of Engel and West (2005) might suffer from the small-sample bias and they adopt the bootstrap method to re-evaluate the causal relationships between exchange rates and fundamentals with the vector autoregressive model. Their bootstrap test results show that the Granger non-causality from exchange rates to the observable fundamentals is not as significant as the existing evidence based on the asymptotic distribution in all sample periods. On the contrary, Bahmani-Oskooee et al. (2015) report cointegration in all 6 multivariate cases, and 20 out of 24 bivariate cases by using the ARDL approach and their Granger non-causality tests report strong evidence that fundamentals help predict exchange rates in both the short run and long run, but not vice versa. There is only weak evidence of unidirectional Granger causality running from exchange rates to market

[18] The empirical results for the multivariate panel model with lag order equals to 1, 2, and 3 are available from the authors upon request.

Table 8
Comparisons with previous studies.

Studies	Methodology	$H_0 : s_t \nRightarrow f_t$	$H_0 : f_t \nRightarrow s_t$
Engel and West (2005)	Johansen's (1991) cointegration test and linear Granger causality test	Rejected. Exchange rates help forecast fundamentals.	In general, no evidence of unidirectional causality running from fundamentals to exchange rates
Bahmani-Oskooee et al. (2015)	Pesaran et al.'s (2001) ARDL cointegration test and linear Granger causality test	Weakly rejected. There is only weak evidence of unidirectional causality running from exchange rates to market fundamentals.	Strongly rejected. The market fundamentals are helpful to forecasting nominal exchange rates
Ko and Ogaki (2015)	Bivariate differenced VAR model with bootstrap Granger causality test	Weakly rejected. the Granger causality from exchange rates to the observable fundamentals is not as significant as the existing evidence	Not available
Evans (2010)	Microstructure approach	Not available	20–30% of the variance in excess currency returns over one- and two-month horizons can be linked back to developments in the macroeconomy
This study	Emirmahmutoglu and Kose's (2011) bootstrap panel Granger non-causality test	Rejected. The exchange rates help predict the money supply, the interest differential and the monetary fundamentals	Strongly rejected. The market fundamentals are helpful to forecasting nominal exchange rates

fundamentals.

Based on the bootstrap panel Granger non-causality test proposed by Emirmahmutoglu and Kose (2011), the empirical results of this study provide strong evidences of unidirectional Granger causality running from the market fundamentals to nominal exchange rates, which are in line with the findings of Bahmani-Oskooee et al. (2015), but do not agree with Engel and West (2005). These results suggest that the monetary approach to exchange rate determination does provide a useful explanation of the behavior of exchange rates. The empirical results of this study also show that nominal exchange rates are helpful to forecasting economic fundamentals of money supply, the interest differential, the inflation differential and the monetary fundamental, respectively. The findings of the Granger causality from exchange rates to fundamentals are in support of Engel and West's (2005) present-value model of exchange rate determination. However, our results are not in line with Ko and Ogaki (2015). The key points for making our empirical results reliable compared to those of Ko and Ogaki (2015) are based on the following two reasons. First, the study of Ko and Ogaki (2015) has the problem of omission of the cross-sectional dependence bias because they examine the causal relationship from exchange rates to fundamentals merely based on the bivariate time series model.[19] We believe that the omission of the cross-sectional dependence of the data indeed has negative impact on detecting the causal relations from exchange rates to fundamentals. This partly explains why they cannot find evidence of Granger causality from exchange rates to fundamentals. This paper alleviates the problem of omission of the cross-sectional dependence by using the bootstrap panel data approach on detecting causal relations. However, bootstrap is not a panacea for modelling the cross-sectional dependence. Unless one knows that nature of the cross-sectional dependence (e.g., exchange rates are automatically cross-sectionally dependent due to the common numeraire, the US dollar), it is difficult to claim that we have modeled the cross-sectional dependence. The bootstrap offers a general solution, but it is more interesting to know the exact or particular source of the cross-sectional dependence.[20]

Second, Ko and Ogaki (2015) implicitly assume that the variables are required to be covariance stationary. Hence, they take the first-difference of the data before conducting the linear Granger non-causality test. The cost of taking the first-difference of the data in testing for the Granger non-causality is that the information regarding the long-run equilibrium relationship inherent in the level data is wasted if it really exists (Engle and Granger, 1987). It in turn results in a possibility of statistical bias in testing for the Granger non-causality. On the contrary, we use the level data combined with the panel bootstrap Granger non-causality test. The advantages of our approach are that we take account of the cross-sectional dependence of the panel data and will not lose any important information in testing for Granger non-causality between exchange rates and fundamentals by using the level data.

Third, a new literature has emerged documenting a strong link between spot rate dynamics and order flows – the transaction flows arising from trades between counterparties in the foreign exchange market. Evans (2010, P. 59) mentions that "In short, an arm's-length observer of the literature might well conclude that the apparent disconnect between spot rates and the macroeconomy is matched by the disconnect between the traditional and microstructure approaches to exchange-rate modelling." The most innovative contribution of Evans' (2010) study is that he propose a theoretical model of exchange-rate determination that bridges the gap between existing microstructure and traditional models. "In particular, the model shows that the order flow generated by trades between dealers and agents can convey information to dealers about the current state of the macroeconomy which they then use to revise their spot exchange rate quotes. Thus, the high frequency behavior of spot exchange rates reflects the flow of new information reaching dealers concerning the slowly evolving state of the macroeconomy, rather than the effects of shocks that drive rapidly changing macroeconomic conditions (Evans, 2010, p. 59, p. 59)." His empirical results indicate that between 20 and 30% of the variance in excess currency returns over one- and two-month horizons can be linked back to developments in the macroeconomy. Evans (2010) provides a straightforward solution to the exchange-rate disconnect puzzle. Readers are referred to Evans and Rime (2012) for a survey on micro approaches to foreign exchange determination or surf Professor Evans' website for a series of studies.[21] This study, based on the traditional model, examines the exchange-rate disconnect puzzle by testing the Granger non-causality between exchange rate and fundamentals, which is proposed by Engel and West (2005). An important empirical implication of the Engel and West (2005) present value model is that current exchange

[19] Basher and Westerlund (2009, p. 506) also stress that "Thus, the empirical performance of the monetary model on an individual country-by-country basis has not been very convincing. But as Taylor and Taylor (2004) point out, time series results of this kind should not be taken too seriously, as the failure to reject the null hypothesis of no cointegration is more likely to reflect the low power of the tests employed rather than the failure of the monetary model."

[20] We thank an anonymous referee for pointing this to us.

[21] Professor Martin D. D. Evans' website: http://faculty.georgetown.edu/evansm1/Home%20page.htm.

rates are helpful for predicting the future fundamentals in terms of Granger's definition. Our empirical evidences show that exchange rates Granger-case the fundamentals, supporting the view that exchange rates are determined as the present value that depends in part on observed fundamentals. The difference between Evans (2010) and this study is that, the former is based on microstructure approach and the latter is based on the traditional model, respectively. Both studies provide empirical evidences of the exchange-rate disconnect puzzle.

5. Concluding remarks

This paper is directed towards revisiting the direction of the Granger non-causality between exchange rates and observed fundamentals implied by the present value model, proposed by Engel and West (2005), under the framework of the monetary model. Compared to previous studies, the key contribution of this paper to the literature is the application of a state-of-the-art Granger non-causality technique that has recently been developed by Emirmahmutoglu and Kose (2011), and it is based on the estimation of the panel model with the bootstrap critical values.

Some interesting conclusions emerge from this empirical study. First, the null hypothesis of no cross-sectional dependence across the members of the panel is strongly rejected, indicating that the bootstrap critical value is required in conducting the panel Granger non-causality test. Therefore, we apply the bootstrap method to generate the empirical distributions of the Fisher test. Second, the test results show that the null hypothesis of Granger non-causality running from economic fundamentals to exchange rates is significantly rejected, implying that the monetary approach to exchange rate determination does provide a useful explanation of the behavior of exchange rates. Therefore, researchers can still rest upon the monetary model as a useful benchmark to understand the evolution of exchange rates in the long run. Third, nominal exchange rates Granger-case the fundamentals such as the money supply, the interest rate differential and the monetary fundamental, by using the bootstrap Granger non-causality tests. A finding of the Granger causality running from exchange rates to observed fundamentals is supportive of a view that exchange rates are determined as the present value that depends in part on observed fundamentals as shown in Engel and West (2005). According to Engel and West (2005, p. 487), "Surely much of the short-run fluctuation in exchange rate changes is driven by changes in expectations about the future. If the models are good approximations and expectations reflect information about future fundamentals, the exchange rate changes will likely be useful in forecasting these fundamentals. So these models suggest that exchange rates Granger-cause the fundamentals." Our empirical results show that the exchange rates Granger-cause the fundamentals, supporting the view that exchange rates are determined as the present value that depends in part on observed fundamentals. Hence, like the stock price and interest rate, the exchange rates do reflect the forward-looking behavior and thus serve as a good predictor of macroeconomic time series such real GDP.[22] However, as pointed by Yuan (2011), observed fundamentals may be important determinants of exchange rates, there may be some other factors, like microstructure effects (Evans and Lyons, 2002) and unobservable trend components (Sarno and Taylor, 2001), driving the exchange rates that current asset-pricing models have not yet captured.

It is of interest to note that, by using the panel cointegration approach, Cerra and Saxena (2010) find strong evidence for cointegration between nominal exchange rates and monetary fundamentals. They also find that the fundamentals-based models are very successful in beating naïve random walk in out-of-sample prediction. Wu and Wang (2013) also show that the Taylor-rule-based fundamental is the best among four different fundamental-based models (i.e., the Taylor-rule-based fundamental, the flexible-price monetary model, the purchasing power parity model, and the uncovered interest parity model) in out-of-sample predictions. In this study, we do not devote to evaluating the out-of-sample forecasting performance of the monetary model by using the panel data approach. This is beyond the scope of this paper. We leave this as a research avenue in the future.

[22] We owe this to an anonymous referee.

Acknowledgements

We would like to thank the editor, Professor Paresh Kumar Narayan, and two anonymous referees of this journal for helpful comments and suggestions. We thank Professor Charles Engel for making his data publicly available in his personal homepage. We also thank An-Chi Wu for her research assistance. The usual disclaimer applies.

Appendix A. Supplementary data

Supplementary data to this article can be found online at https://doi.org/10.1016/j.econmod.2018.09.021.

References

Bahmani-Oskooee, M., Hegerty, S., Tanku, A., 2010. The black-market exchange rate versus the official rate: which rate fosters the adjustment speed in the monetarist model? Manch. Sch. 78 (6), 725–738.
Bahmani-Oskooee, M., Hosny, A., Kishor, N.K., 2015. The exchange rate disconnect puzzle revisited. Int. J. Finance Econ. 20, 126–137.
Bai, J., Kao, C., 2006. On the estimation and inference of a panel cointegration method with cross-sectional dependence. In: Baltagi, B.H. (Ed.), Panel Data Econometrics: Theoretical Contributions and Empirical Applications. Elsevier Science and Technology.
Balke, N.S., Ma, J., Wohar, M.E., 2013. The contribution of economic fundamentals to movements in exchange rates. J. Int. Econ. 90, 1–16.
Basher, S.A., Westerlund, J., 2009. Panel cointegration and the monetary exchange rate model. Econ. Modell. 26, 506–513.
Baxter, M., Stockman, A., 1989. Business cycles and the exchange rate regime. J. Monetary Econ. 23 (3), 377–400.
Beckmann, J., Belke, A., Kühl, M., 2011. The dollar-euro exchange rate and macroeconomic fundamentals: a time-varying coefficient approach. Rev. World Econ. 147, 11–40.
Beckmann, J., Belke, A., Dobnik, F., 2012. Cross-section dependence and the monetary exchange rate model – a panel analysis. N. Am. J. Econ. Finance 23 (1), 38–53.
Bekiros, S.D., 2014. Exchange rates and fundamentals: Co-movement, long-run relationships and short-run dynamics. J. Bank. Finance 39, 117–134.
Bilson, J., 1978a. Rational expectations and the exchange rate. In: Frenkel, J., Johnson, H. (Eds.), The Economics of Exchange Rates. Addison-Wesley, Reading, Massachusetts, pp. 75–96.
Bilson, J., 1978b. The monetary approach to the exchange rate: some empirical evidence. IMF Staff Pap. 25 (1), 48–75.
Breitung, J., 2005. A parametric approach to the estimation of cointegration vectors in panel data. Econom. Rev. 24, 151–173.
Breusch, T., Pagan, A., 1980. The Lagrange multiplier test and its application to model specifications in econometrics. Rev. Econ. Stud. 47, 239–253.
Campbell, J.Y., Shiller, R.J., 1988a. The dividend-price ratio and the expectations of future dividends and discount factors. Rev. Financ. Stud. 1, 195–228.
Campbell, J.Y., Shiller, R.J., 1988b. Stock prices, earnings and expected dividends. J. Finance 43, 661–676.
Cerra, V., Saxena, S.C., 2010. The monetary model strikes back: evidence from the world. J. Int. Econ. 81, 184–196.
Chen, S.-S., Chou, Y.-H., 2015. Revisiting the relationship between exchange rates and fundamentals. J. Macroecon. 46, 1–22.
Chang, M.-J., Su, C.-Y., 2014. The dynamic relationship between exchange rates and macroeconomic fundamentals: evidence from Pacific Rim countries. J. Int. Financ. Market. Inst. Money 30, 220–246.
Cheung, Y.-W., Chinn, M.D., Pascual, A.G., 2005. Empirical exchange rate models of the nineties: are any fit to survive? J. Int. Money Finance 24 (7), 1150–1175.
Choi, I., 2001. Unit root test for panel data. J. Int. Money Bank. 20, 249–272.
Civcir, I., 2004. The long-run validity of the monetary exchange rate model for a high inflation country and misalignment: the case of Turkey. Emerg. Market. Finance Trade 40 (4), 84–100.
Clarida, R., Sarno, L., Taylor, M., Valente, G., 2003. The out-of-sample success of term structure models as exchange rate predictors: a step beyond. J. Int. Econ. 60, 61–83.
Cushman, D.O., 2000. The failure of the monetary exchange rate model for the canadian-U.S. dollar. Can. J. Econ. 84, 1161–1176.
Dabrowski, M.A., Papiez, M., Smiech, S., 2014. Exchange rates and monetary fundamentals in CEE countries: evidence from a panel approach. J. Macroecon. 41, 148–159.

Demetrescu, M., Hanck, C., 2012. A simple nonstationary-volatility robust panel unit root test. Econ. Lett. 117 (2), 10–13.
Diks, C., Panchenko, V., 2005. A note on the Hiemstra-Jones test for Granger non-causality. Stud. Nonlinear Dynam. Econom. 9 (art. 4).
Diks, C., Panchenko, V., 2006. A new statistic and practical guidelines for nonparametric Granger causality testing. J. Econ. Dynam. Contr. 30, 1647–1669.
Dornbusch, R., 1976. Expectations and exchange rate dynamics. J. Polit. Econ. 84 (6), 1161–1176.
Emirmahmutoglu, F., Kose, N., 2011. Testing for Granger causality in heterogeneous mixed panels. Econ. Modell. 28, 870–876.
Engel, C., 1994. Can the Markov switching model forecast exchange rates? J. Int. Econ. 36, 151–165.
Engel, C., 1996. The forward discount anomaly and the risk premium: a survey of recent evidence. J. Empir. Finance 3, 123–191.
Engel, C., Hamilton, J., 1990. Long swings in the dollar: are they in the data and do market know it? Am. Econ. Rev. 80, 687–713.
Engel, C., West, K.D., 2004. Accounting for exchange-rate variability in present-value models when the discount factor is near 1. AEA Pap. Proc. 94 (2), 119–125.
Engel, C., West, K.D., 2005. Exchange rates and fundamentals. J. Polit. Econ. 113 (3), 485–517.
Engel, C., West, K.D., 2006. Taylor rules and the Deutschemark-Dollar real exchange rate. J. Money Credit Bank. 38, 1175–1194.
Engel, C., Mark, N.C., West, K.D., 2007. Exchange rate models are not as bad as you think. In: Acemoglu, Daron, Rogoff, Kenneth, Woodford, Michael (Eds.), NBER Macroeconomics Annual 2007. MIT Press, Cambridge, MA, pp. 381–411.
Engle, R.F., Granger, C.W.J., 1987. Cointegration and error correction: representation, estimation and testing. Econometrica 55, 251–276.
Evans, M.D.D., 2010. Order flows and the exchange rate disconnect puzzle. J. Int. Econ. 80, 58–71.
Evans, M.D.D., Lyons, R.K., 2002. Order flow and exchange rate dynamics. J. Polit. Econ. 110, 170–180.
Evans, M.D.D., Rime, D., 2012. Micro Approaches to Foreign Exchange Determination, Handbook of Exchange Rates, pp. 73–110 (Chapter 3).
Fisher, R.A., 1932. Statistical Methods for Research Workers, fourth ed. Oliver and Boyd, Edinburgh.
Flood, R., Rose, A., 1995. Fixing exchange rates: a virtual quest for fundamentals. J. Monetary Econ. 36, 3–37.
Francis, B., Hasan, I., Lothian, J., 2001. The monetary approach to exchange rates and the behavior of the Canadian dollar over the long-run. Appl. Financ. Econ. 11, 475–481.
Frankel, J.A., 1979. On the mark: a theory of exchange rates based on real interest differentials. Am. Econ. Rev. 69 (4), 610–622.
Frankel, J.A., 1984. Tests of monetary and portfolio balance models of exchange rate determination. In: Bilson, John F.O., Marston, Richard C. (Eds.), Chapter in NBER Book Exchange Rate Theory and Practice, pp. 239–260.
Frenkel, J.A., 1976. A monetary approach to the exchange rate: doctrinal aspects and empirical evidence. Scand. J. Econ. 78, 200–224.
Frenkel, M., Koske, I., 2004. How well can monetary factors explain the exchange rate of the euro? Atl. Econ. J. 32 (3), 232–243.
Frömmel, M., MacDonald, R., Menkhoff, L., 2005a. Markov switching regimes in a monetary exchange rate model. Econ. Modell. 22, 485–502.
Frömmel, M., MacDonald, R., Menkhoff, L., 2005b. Do fundamentals matter for the D-Mark/Euro-Dollar? A regime switching approach. Global Finance J. 15, 321–335.
Gonzalo, J., Ng, S., 2001. A systematic framework for analyzing the dynamic effects of permanent and transitory shocks. J. Econ. Dynam. Contr. 25, 1527–1546.
Granger, C.W.J., 2003. Some aspects of causal relationships. J. Econom. 112, 69–71.
Groen, J.J.J., 2000. The monetary exchange rate model as a long phenomenon. J. Int. Econ. 52, 299–319.
Groen, J.J.J., 2002. Cointegration and the monetary exchange rate model revisited. Oxf. Bull. Econ. Stat. 64, 361–380.
Grauwe, P.D., Vansteenkiste, I., 2007. Exchange rates and fundamentals: a non-linear relationship? Int. J. Finance Econ. 12, 37–54.
Herwartz, H., Siedenburg, F., 2008. Homogeneous panel unit root tests under cross sectional dependence: finite sample modifications and the wild bootstrap. Comput. Stat. Data Anal. 53 (1), 137–150.
Herwartz, H., Maxand, S., Raters, F.H.C., Walle, Y.M., 2018. Panel unit root tests for heteroskedastic panels. STATA J. 18 (1), 184–196.
Herwartz, H., Maxand, S., Walle, Y.M., 2017. Heteroskedasticity-robust Unit Root Testing for Trending Panels. Center for European, Governance and Economic Development Research. discussion papers 314.
Im, K., Pesaran, M.H., Shin, Y., 2003. Testing for unit roots in heterogeneous panels. J. Econom. 115, 53–74.
Islam, M.F., Hasan, M.S., 2006. The monetary model of the Dollar-Yen exchange rate determination: a cointegration approach. Int. J. Bus. Econ. 5 (2), 129–145.
Johansen, S., 1991. Estimation and hypothesis testing of cointegration vectors in Gaussian vector autoregressive models. Econometrica 59, 1551–1580.
Kar, M., Nazlioglu, S., Agir, H., 2011. Financial development and economic growth nexus in the MEAN countries: bootstrap panel Granger causality analysis. Econ. Modell. 28, 685–693.
Kilian, L., Taylor, M.P., 2003. Why is it so difficult to beat the random walk forecast of exchange rates? J. Int. Econ. 60, 85–107.
Kim, B.-H., Min, H.,-G., Moh, Y.-K., 2010. Nonlinear dynamics in exchange rate deviations from the monetary fundamentals: an empirical study. Econ. Modell. 27, 1167–1177.
Klaassen, F., 2005. Long swings in exchange rates: are they really in the data? J. Bus. Econ. Stat. 23 (1), 87–95.
Ko, H.-H., Ogaki, M., 2015. Granger causality from exchange rates to fundamentals: what does the bootstrap test show us? Int. Rev. Econ. Finance 38, 198–206.
Kónya, L., 2006. Exports and growth: Granger causality analysis on OECD countries with a panel data approach. Econ. Modell. 23 (6), 978–992.
Laganá, G., Sgro, P.M., 2007. The exchange rate disconnected puzzle: a resolution? Asia-Pac. J. Account. Econ. 14, 43–68.
Maddala, G.S., Wu, S., 1999. A comparative study of unit root tests with panel data and a new simple test. Oxf. Bull. Econ. Stat. 61, 631–652.
Mark, N.C., 1995. Exchange rates and fundamentals: evidence on long-horizon predictability. Am. Econ. Rev. 85 (1), 201–218.
Mark, N.C., Sul, D., 2001. Nominal exchange rates and monetary fundamentals: evidence from a small post-Bretton Woods panel. J. Int. Econ. 53, 29–52.
Meese, R.A., Rogoff, K.S., 1983a. Empirical exchange rate models of the seventies: do they fit out of sample? J. Int. Econ. 14, 3–24.
Meese, R.A., Rogoff, K.S., 1983b. The out of sample failure of empirical exchange ratemodels: sampling error or misspecification? In: Frenkel, J.A. (Ed.), Exchange Rates and International Macroeconomics. University of Chicago Press, Chicago, pp. 67–112 (for NBER).
Mussa, M., 1976. The exchange rate, the balance of payments and monetary and fiscal policy under a regime of controlled floating. Scand. J. Econ. 78, 229–248.
Michael, P., Nobay, R., Peel, D., 1997. Transactions costs and nonlinear adjustment in real exchange rates: an empirical investigation. J. Polit. Econ. 105, 862–879.
Neely, C., Sarno, L., 2002. How well do monetary fundamentals forecast exchange rates? Fed. Reserv. Bank St. Louis 84, 51–74.
Obstfeld, M., Rogoff, K., 2001. The six major puzzles in international macroeconomics: is there a common cause? NBER Macroecon. Annu. 2000 (15), 339–390.
Palm, F.C., Smeekes, S., Urbain, J.-P., 2011. Cross-sectional dependence robust block bootstrap panel unit root tests. J. Econom. 163, 85–104.
Panopoulou, E., Pittis, N., 2004. A comparison of autoregressive distributed lag and dynamic OLS cointegration estimators in the case of serially correlated cointegration error. Econom. J. 7, 585–617.
Pesaran, M.H., 2004. General Diagnostic Tests for Cross Section Dependence in Panels. CESifo Working Paper 1229, IZA Discussion Paper 1240.
Pesaran, M.H., Shin, Y., Smith, R.J., 2001. Bounds testing approaches to the analysis of level relationships. J. Appl. Econom. 16 (3), 289–326.
Pesaran, M.H., Ullah, A., Yamagata, T., 2008. A bias-adjusted LM test of error cross-section independence. Econom. J. 11, 105–127.
Pilbeam, K., 1995. Exchange rate models and exchange rate expectations: an empirical investigation. Appl. Econ. 27, 1009–1015.
Rapach, D.E., Wohar, M.E., 2002. Testing the monetary model of exchange rate determination: new evidence from a century of data. J. Int. Econ. 58, 359–385.
Rapach, D.E., Wohar, M.E., 2004. Testing the monetary model of exchange rate determination: a closer look at panels. J. Int. Money Finance 23, 867–895.
Rogoff, K., 1996. The purchasing power parity puzzle. J. Econ. Lit. 34, 647–668.
Sarantis, N., 1994. The monetary exchange rate model in the long run: an empirical investigation. Weltwirtschaftliches Archiv 130, 698–711.
Sarno, L., 2005. Towards a solution to the puzzles in exchange rate economics: where do we stand? Can. J. Econ. 38 (3), 673–708.
Sarno, L., Sojli, E., 2009. The feeble link between exchange rates and fundamentals: can we blame the discount factor? J. Money Credit Bank. 41, 437–442.
Sarno, L., Taylor, M.P., 2001. *The* Microstructure of the Foreign Exchange Market, Princeton Studies in International Economics. Princeton University.
Shen, C.-H., Chen, S.-W., 2004. Long swing in appreciation and short swing in depreciation and does the market not know it — the case of Taiwan. Int. Econ. J. 18 (2), 195–213.
Sims, C.A., Stock, J.H., Watson, M.W., 1990. Inference in linear time series models with some unit roots. Econometrica 58, 113–144.
Smeekes, S., Urbain, J.-P., 2014. On the applicability of the sieve bootstrap in time series panels. Oxf. Bull. Econ. Stat. 76 (1), 139–151.
Westerlund, J., Smeekes, S., 2018. Robust block bootstrap panel predictability tests. Econom. Rev. (In Press).
Stine, R.A., 1987. Estimating properties of autoregressive forecasts. J. Am. Stat. Assoc. 82, 1072–1078.
Swamy, P.A.V.B., 1970. Efficient inference in a random coefficient regression model. Econometrica 38, 311–323.
Tawadros, G.B., 2001. The predictive power of the monetary model of exchange rate determination. Appl. Financ. Econ. 11 (3), 279–286.
Tawadros, G.B., 2017. Revisiting the exchange rate disconnect puzzle. Appl. Econ. 49, 3645–3668.
Taylor, M., Peel, D., 2000. Non-linear adjustment, long-term equilibrium and exchange rate fundamentals. J. Int. Money Finance 19, 33–53.
Taylor, M., Peel, D., Sarno, L., 2001. Non-linear adjustment in real exchange rates: towards a solution to the purchasing power parity puzzles. Int. Econ. Rev. 42, 1015–1042.
Taylor, A.M., Taylor, M.P., 2004. The purchasing power parity debate. J. Econ. Perspect. 18, 135–158.
Toda, H.Y., Phillips, P.C.B., 1993. Vector autoregression and causality. Econometrica 59, 229–255.
Toda, H.Y., Yamamoto, T., 1995. Statistical inference in vector autoregressions with possibly integrated processes. J. Econom. 66, 225–250.
Westerlund, J., 2006. Testing for panel cointegration with multiple structural breaks. Oxf. Bull. Econ. Stat. 68, 101–132.
Westerlund, J., 2014. Heteroscedasticity robust panel unit root tests. J. Bus. Econ. Stat. 32, 112–135.
Westerlund, J., 2015. The effect of recursive detrending on panel unit root tests. J. Econom. 185, 453–467.

Wu, J.-L., Hu, Y.-H., 2009. New evidence on nominal exchange rate predictability. J. Int. Money Finance 28, 1045–1063.
Wu, J.-L., Wang, Y.-C., 2013. Fundamentals, forecast combinations and nominal exchange-rate predictability. Int. Rev. Econ. Finance 25, 129–145.
Yamada, H., Toda, H.Y., 1998. Inference in possibly integrated vector autoregressive models: some finite sample evidence. J. Econom. 86, 55–95.
Yuan, C., 2011. The exchange rate and macroeconomic determinants: time-varying transitional dynamics. N. Am. J. Econ. Finance 22, 197–220.
Zellner, A., 1962. An efficient method of estimating seemingly unrelated regressions and tests for aggregation bias. J. Am. Stat. Assoc. 57, 348–368.