Economic Modelling 40 (2014) 42–51

Contents lists available at ScienceDirect

Economic Modelling

journal homepage: www.elsevier.com/locate/ecmod

Money demand instability and real exchange rate persistence in the monetary model of USD–JPY exchange rate

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ARTICLE INFO

Article history:
Accepted 13 March 2014
Available online 12 April 2014

Keywords:
Cointegration
Exchange rates
Monetary model
Weak exogeneity

ABSTRACT

This paper proposes a hybrid monetary model of the dollar–yen exchange rate that takes into account factors affecting the conventional monetary model's building blocks. In particular, the hybrid monetary model is based on the incorporation of real stock prices to enhance money demand stability and also, productivity differential, relative government spending, and real oil price to explain real exchange rate persistence. By using quarterly data over a period of high international capital mobility and volatility (1980:01–2009:04), the results show that the proposed hybrid model provides a coherent long-run relation to explain the dollar–yen exchange rate as opposed to the conventional monetary model.

1. Introduction

Since the collapse of the Bretton Woods fixed exchange rate system in 1971, much attention has been paid towards finding a meaningful explanation of exchange rates. A wide range of models have been proposed to understand movements in the exchange rate, one of which is the monetary model (see Bilson, 1978; Frankel, 1979). Despite its rigorous theoretical underpinnings by linking the nominal exchange rate to its monetary fundamentals (e.g., money, income, and interest rates), the resulting reduced form has had limited empirical success until now.

For example, although MacDonald and Taylor (1994) provided evidence of a long-run relation between monetary fundamentals and nominal exchange rates, the signs and magnitudes of estimated coefficients did not support the related monetary theories. Groen (2000), and Mark and Sul (2001) among others also found some evidence in a panel context, but this was under the assumption of a high order of heterogeneity across the country models. Similarly, Rapach and Wohar (2002) found some support for the theory using long time series, but this was related to different exchange rates and macro regimes, with some evolution in the composition of products in price indices. Taylor and Peel (2000) applied nonlinear methods to model a nominal exchange rate and monetary fundamentals (relative money supply and relative income), but such results are often sensitive to a small number of observations and become less robust as the sample evolves. Frömmel et al. (2005) estimated the real interest differential (RID) model of Frankel (1979) applying the Markov switching approach. However, the model was shown to relate to only one regime.

Furthermore, the empirical failure of this model has been specifically found in regard of the US dollar–Japanese yen exchange rate. The evolution of this exchange rate has been much debated over the recent years with no consensus over the factors that drive the dynamics. For instance, Caporale and Pittis (2001) were unable to find a stable relation based on a monetary model of this exchange rate. Chinn and Moore (2011) also failed to find a long-run relation between the nominal dollar–yen exchange rate and its monetary fundamentals (money, industrial production, and interest rate differentials) even when they included cumulative order flow as opposed to the dollar–euro exchange rate. By contrast, MacDonald and Nagayasu (1998) only found that a simplified version of the RID model of Frankel (1979), that excluded the money demand functions, held for the yen–dollar exchange rate for the period 1975:Q3–1994:Q3. Tellingly, in a recent paper, Obstfeld (2009, p.1) comments that 'the determinants of the yen's short- and even longer-term movements remain mysterious in light of the development of Japan's macro economy'.

A possible explanation for the empirical failure of the dollar–yen exchange rate monetary model is perhaps the breakdown of its underlying building blocks; that is, stable money demand and purchasing power parity (PPP). Indeed, Hendry and Ericsson (1991) found that the conventional money demand equation for the US was not stable. Whereas, Friedman (1988) and McCornac (1991) confirmed the need for real stock prices to stabilise money demand equations using data from the United States and Japan, respectively.[1] Sarno and Taylor (2002), on the other hand, found little support for the conventional notion of PPP by surveying a range of empirical studies. This corresponds well with the classic findings of

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[1] The conclusion of the seminal article of Friedman (1988) has also been confirmed by Choudhry (1996) for the US and Canada and Caruso (2001) for Japan, the UK, Switzerland and Italy, as well as for a panel of 25 (19 industrial and 6 developing) countries.

http://dx.doi.org/10.1016/j.econmod.2014.03.019
0264-9993/

Balassa (1964) and Samuelson (1964), which indicate that persistent deviations from PPP arise from productivity differentials. Chinn (1997, 2000), and Wang and Dunne (2003) among others showed that fluctuations in the nominal and real dollar–yen exchange rate are due to the impact of differentials in productivity and government expenditure along with real oil prices.

This paper contributes to the existing literature by proposing a hybrid monetary model of the dollar–yen exchange rate that takes into account the breakdown of the aforementioned building blocks. That is, the proposed model captures both the monetary and the real aspects of the economy, thereby circumventing some of the potential pitfalls associated with earlier studies. More specifically, we examine the empirical performance of the standard RID model, developed by Frankel (1979), against this proposed hybrid version by employing the Johansen (1995) methodology and quarterly data from 1980:01 to 2009:04, a period characterised by high international capital mobility and volatility.

The RID model has been widely used as it combines aspects of the sticky-price approach with the flexible-price one. Furthermore, this variant of the monetary approach is chosen because it is a realistic description when variation in the inflation differential is moderate as is the case between the US and Japan over the period under examination.[2] Particularly, the theory underlines the role of expectations in different inflationary environments and the associated rapid adjustment in capital markets. The hybrid version, by contrast, is devised by using domestic and foreign money demand equations based on broader asset classes and also accounting for the factors that cause PPP to fail. That is, we incorporate real stock prices in the money demand equations,[3] while we use the productivity differential, relative government spending, and real oil price to explain the persistence in the real dollar–yen exchange rate.

The paper is organised as follows. Section 2 provides the theoretical framework for the exchange rate monetary model; Section 3 outlines the econometric technique used and describes the data; Section 4 explains the empirical results and the analysis; and finally Section 5 concludes.

2. Theoretical framework

The monetary model of the exchange rate is based on the assumptions that money demand equations are stable and that PPP holds. In this paper, we consider two forms of this model and place them under econometric scrutiny. The first is the RID model developed by Frankel (1979) and the second is a hybrid monetary model, proposed herein, that takes into account factors affecting the stability of the respective money demand equations and the validity of PPP.

In Frankel's (1979) RID model, the features of the fixed- and flexible-price models are amalgamated by incorporating short-term interest rates to capture the stance of monetary policy. In particular, the model asserts that the expected rate of depreciation of the exchange rate is a function of the gap between the current spot rate and the long-run equilibrium rate, as well as the expected long-run inflation differential between the domestic and foreign countries (see Pilbeam, 2013, Chapter 7); that is:

$$E(\Delta e) = \theta(\bar{e}-e) + \left(\Delta p^{e}-\Delta p^{*e}\right), \tag{1}$$

where θ is the speed of adjustment towards the equilibrium level and, Δp^e and Δp^{*e} denote the domestic and foreign expected long-run inflation rates, respectively. Note that throughout the paper all variables are expressed in natural logs (except interest rates), bars denote equilibrium values, and the asterisk denotes the foreign country (Japan) and the domestic country is the United States. It follows that Eq. (1) highlights that in the short-run the spot exchange rate e is expected to return to its long-run equilibrium value $\bar{e}$ at a rate equal to θ. However, in the long-run (since $\bar{e} = e$), changes in the exchange rate will be proportional to the expected long-run inflation differential ($\Delta p^e - \Delta p^{*}{}^{e}$).

Assuming the uncovered interest parity (UIP) condition, $E(\Delta e) = i - i^*$, that postulates domestic and foreign bonds are perfect substitutes, then combining such a condition with Eq. (1) and rearranging for the spot exchange rate, we obtain:

$$e = \bar{e}-\frac{1}{\theta}\left[\left(i-\Delta p^{e}\right)-\left(i^{*}-\Delta p^{e*}\right)\right], \tag{2}$$

where i and i^* are defined as the domestic and foreign interest rates, respectively. Furthermore, conventional domestic and foreign money demand equations are given as follows:

$$m-p = a_1 y-a_2 i, \tag{3}$$

$$m^{*}-p^{*} = a_1 y^{*}-a_2 i^{*}, \tag{4}$$

where m (m^*), p (p^*), and y (y^*) are respectively domestic (foreign) money supply, price level, and real income. For simplicity, the income elasticity of money demand, a_1, and the interest rate semi-elasticity of money demand, a_2, are assumed to be identical across both domestic and foreign countries. Also, it is assumed that PPP holds in the long-run:

$$\bar{e} = \bar{p}-\overline{p^{*}}. \tag{5}$$

By extracting the expressions of relative prices in Eq. (5) from Eqs. (3) and (4) along with the view that $\left(\bar{i}-\bar{i}^{*}\right) = (\Delta p^{e}-\Delta p^{*e})$ in the long-run (since $\bar{e} = e$), the following is obtained with bars denoting equilibrium values:

$$\bar{e} = \left(\overline{m}-\overline{m}^{*}\right)-a_1\left(\bar{y}-\bar{y}^{*}\right) + a_2\left(\Delta p^{e}-\Delta p^{*e}\right). \tag{6}$$

By substituting Eq. (6) into Eq. (2), we obtain:

$$\begin{aligned} e = {} & \left(\overline{m}-\overline{m}^{*}\right)-a_1\left(\bar{y}-\bar{y}^{*}\right) \\ & + a_2\left(\Delta p^{e}-\Delta p^{*e}\right)-\frac{1}{\theta}\left[\left(i-\Delta p^{e}\right)-\left(i^{*}-\Delta p^{*e}\right)\right]. \end{aligned} \tag{7}$$

Frankel (1979) argued that it is common practice to estimate this equation empirically on the basis that short-term interest rates represent real interest rates (i.e., liquidity effects of monetary policy) and long-term interest rates capture the long-run expected inflation rates (see also MacDonald, 2007, Chapter 6). Thus, the baseline model is in the reduced form written as follows:

$$e_t = \beta_1\left(m_t-m_t^{*}\right) + \beta_2\left(y_t-y_t^{*}\right) + \beta_3\left(i_t^{s}-i_t^{s*}\right) + \beta_4\left(i_t^{l}-i_t^{l*}\right) + \varepsilon_t. \tag{8}$$

Otherwise, the RID model related to Eq. (8) hypothesises that an increase in the domestic money supply relative to the counterpart foreign one increases domestic prices and thus causes a one for one depreciation in the exchange rate ($\beta_1 = 1$). An increase in domestic income or a decline in the expected rate of domestic inflation (proxied by the long-term interest rate) relative to the foreign one raises the demand for money and thus causes an appreciation in the exchange rate ($\beta_2 < 0$, $\beta_4 > 0$). An increase in the domestic nominal interest rate relative to the foreign one induces capital inflows towards the domestic economy and thus causes an appreciation in the exchange rate ($\beta_3 < 0$). For further details the reader is directed to Frankel (1979).

[2] Bernanke (2000) and Taylor (2001) argued that the different inflationary environments in the US and Japan are due to the differences in the monetary policies in the two countries.

[3] Another motivation for incorporating real stock prices in the monetary model via money demand equations is that the financial press and financial market analysts advocate that there exists a relation between stock prices and exchange rates (see, for examples, Caporale et al., 2014; Phylaktis and Ravazzola, 2005).

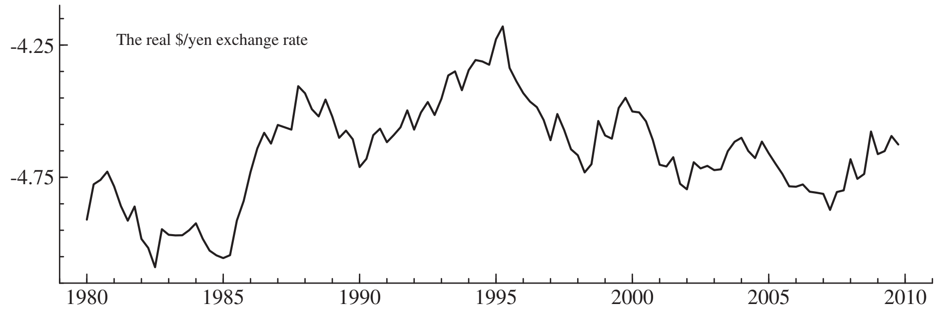

Fig. 1. Behaviour of the real dollar–yen exchange rate for the period 1980:Q1–2009:Q4.

However, Friedman (1988) and subsequently McCornac (1991) and Caruso (2001) among others showed that the stability of the money demand functions used to specify the monetary model, Eqs. (3) and (4), depends on the inclusion of real stock prices. Furthermore, as Chortareas and Kapetanios (2004) pointed out, there is limited support for the conventional notion of PPP for Japan. Indeed, by visual inspection of Fig. 1, we find that the real dollar–yen exchange rate, calculated as the nominal exchange rate adjusted for the domestic and foreign price levels (see Eq. (6a) in Appendix A), does not appear to revert to mean.

Balassa (1964) and Samuelson (1964) attributed the inadequacy of PPP to real economic shocks, in particular, to the unanticipated movement found in the productivity differentials between the traded and non-traded goods sectors across the economies. Financial variables also appear sensitive to the demand shocks associated with government expenditure (Chinn, 2000) and the supply shocks related to real oil prices (Amano and van Norden, 1998). In the context of the yen–dollar exchange rate, Chinn (1997, 2000), and Wang and Dunne (2003) found that real economic factors were responsible for any persistence in the real yen–dollar exchange rate during the post-Bretton Woods period.

Using these factors, we amend Eq. (8) accordingly and term it the hybrid monetary (HM) model which, as derived in Appendix A, takes the following form:

$$e_t = \beta_1\left(m_t - m_t^{*}\right) + \beta_2\left(y_t - y_t^{*}\right) + \beta_3\left(i_t^{s} - i_t^{s*}\right) + \beta_4\left(i_t^{l} - i_t^{l*}\right) + \beta_5\left(s_t - s_t^{*}\right) + \beta_6\left(prod_t^{T} - prod_t^{T*}\right) + \beta_7\left(gs_t - gs_t^{*}\right) + \beta_8\, roil_t + v_t, \tag{9}$$

where s_t stands for the real stock prices, gs_t is the government consumption as a percentage of GDP, $prod_t^T$ is the productivity in the traded sector, and $roil_t$ is the real oil price.

In addition to the coefficient restrictions discussed earlier ($\beta_1 = 1, \beta_2 < 0, \beta_3 < 0, \beta_4 > 0$), the HM model suggests that the sign of the coefficient on real stock prices, β_5, depends on the extent to which the substitution effect (positive) dominates the wealth effect (negative) in the money demand equation. Based on the derivation provided in Appendix A, the sign of the coefficient on the productivity differential depends on the relative competitiveness of the traded goods sector. Specifically, an increase in the productivity of the traded sector relative to the non-traded sector in the domestic economy compared to the foreign one results in a fall in the domestic traded sector's goods prices relative to the foreign counterpart, and then an exchange rate appreciation ($\beta_6 < 0$).

The differential in government expenditure captures differences in demand side shocks (Chinn, 2000). As government expenditure is anticipated to be spent largely on non-tradable goods such as services, an increase in domestic government spending relative to the foreign counterpart should then increase the relative price of domestic non-tradable goods, leading to an exchange rate appreciation ($\beta_7 < 0$). The sign of the coefficient on the real oil price is expected to be negative ($\beta_8 < 0$) because oil price is given in the US dollar and higher real oil price should lead to an appreciation in the dollar (see Amano and van Norden, 1998). That is, the input costs in Japan are highly sensitive to the oil price because Japan is a net importer country and the third largest oil consumer and importer country after the United States and China.[4]

3. The econometric approach and data

3.1. The econometric approach

We employ the Johansen methodology to investigate the long-run equilibrium relations between the variables of the two models, RID and HM. Johansen (1995) formulates an unrestricted VAR model of order p with $(n \times 1)$ endogenous variables, all integrated of order one $(I(1))$, forced by a vector of $(n \times 1)$ independent Gaussian errors, with the following error-correction representation:

$$\Delta X_t = \Pi X_{t-1} + \Gamma_1 \Delta X_{t-1} + \ldots + \Gamma_p \Delta X_{t-(p-1)} + \gamma D_t + \varepsilon_t, \tag{10}$$

where X_t is an $(n \times 1)$ vector of variables; D_t is a vector containing constants, centred seasonal dummies, and impulse dummies; Γ_i $(i = 1, \ldots, p-1)$ are $(n \times n)$ parameter matrices capturing the short-run dynamics among the variables; and Π is an $(n \times n)$ matrix decomposed as $\alpha\beta'$, with matrices α and β dimensioned $(n \times r)$, relating to the speed of adjustment and long-run relations, respectively.

We use the trace test to determine the rank r of Π. Johansen (1995) explains that the test has an optimal sequence starting with the null hypothesis $r = 0$ (no cointegration) against the alternative $r \le 1$ (at least one cointegrating vector) and subsequent further orders of cointegration $r = i$ against the alternative $r \le i + 1$; the sequence stops at $r = i$ when the null cannot be rejected. The test is a likelihood ratio test that can be written in terms of eigenvalues (λ_i) and sample size (T) with

$$\lambda_{trace} = -T\sum_{i=1}^{n}(1-\lambda_i). \tag{11}$$

The results associated with the Johansen test are well-defined when the VAR model is well-specified (Johansen, 1995). The most appropriate lag length for the model is often selected on the basis of information criteria such as the Schwarz Bayesian information criterion (SBIC), the Akaike information criterion (AIC), and the Hannan–Quinn information criterion (HQIC). However, Burke and Hunter (2007) suggest that there can be substantial size distortion of the trace test relative to the null

[4] Japanese oil consumption and imports in 2010 were respectively 23% and 42.7% of the consumption and imports of the United States; figures obtained online from the CIA World Factbook (2011). However, the US is rich in natural resources.

Table 1
ADF unit root test results.

	Level		First difference	
	Constant	Constant and trend	Constant	Constant and trend
e	−1.212 (3)	−2.058 (4)	−5.492 (2)[a]	−5.456 (2)[a]
$m-m^*$	−1.450 (4)	−2.142 (4)	−3.403 (4)[b]	−3.428 (4)
$y-y^*$	−0.402 (3)	−2.250 (3)	−3.762 (2)[a]	−3.807 (2)[b]
i^s-i^{s*}	−2.509 (7)	−2.494 (7)	−4.960 (7)[a]	−4.936 (7)[a]
i^l-i^{l*}	−1.803 (6)	−2.512 (6)	−5.742 (5)[a]	−5.700 (5)[a]
$s-s^*$	−0.433 (0)	−1.618 (0)	−9.925 (0)[a]	−9.616 (0)[a]
$gs-gs^*$	−0.383 (3)	−2.012 (3)	−4.211 (2)[a]	−4.213 (2)[a]
$prod^T-prod^{T*}$	0.567 (5)	−2.562 (5)	−5.671 (4)[a]	−5.666 (4)[a]
$roil$	−1.612 (5)	−1.609 (5)	−6.405 (4)[a]	−6.876 (4)[a]

Note: The 1% and 5% critical values for the ADF test are respectively −3.486 and −2.885 (without trend) and −4.036 and −3.448 (with the trend); the proper lag length, allowing for a maximum of eight lags, is selected on the basis of the general-to-specific approach, represented in parentheses. [a] and [b] indicate significance at the 1% and 5% levels, respectively.

distribution when the selected lag order is sub-optimal.[5] Therefore, we extend the model to include adequate lags to remove any serial correlation in case the lag selected based on information criteria does not capture the dynamics.

As a result of sharp changes as well as differences in monetary policy between the United States and Japan throughout the sample period, we also include impulse dummies that remove the impact of extreme observations relating to 1980:4, 1982:3, 2002:2, and 2008:4. The corresponding known events for the first two dummies relate to the large short-term interest rate fluctuations in the United States and Japan in the late 1970s and early 1980s. Note that the fourth quarter of 1980 also corresponds with the end point of the fiscally liberal 60s and 70s that led to the election of Ronald Reagan as the US President and the Volker reforms at the Federal Reserve. The third dummy corresponds to the monetary expansions (now termed quantitative easing (QE)) adopted by the Bank of Japan from March 2001 to March 2003, while the fourth is due to QE in the United States as a result of the 2007–2008 banking crisis.

Our modelling approach follows Juselius and MacDonald (2004) who consider the joint modelling of the international parity relations between the United States and Japan. More specifically, we examine the RID and HM models econometrically by estimating Eq. (10) using the following variable vectors in their respective levels:

$$X'_{RID,t} = \left[e_t, m_t-m_t^*, y_t-y_t^*, i_t^s-i_t^{s*}, i_t^l-i_t^{l*}\right].$$

$$X'_{HM,t} = \left[e_t, m_t-m_t^*, y_t-y_t^*, i_t^s-i_t^{s*}, i_t^l-i_t^{l*}, s_t-s_t^*, prod_t^T-prod_t^{T*}, gs_t-gs_t^*, roil_t\right].$$

We suggest that by investigating these two variable sets, we might be able to determine the key factors that identify the long-run monetary model of the dollar–yen exchange rate and explain the short-run behaviour of the different systems. Using series that are $I(1)$, we can observe an exchange rate equation based on the model in question by finding a cointegrating relation and showing via a likelihood ratio test that this variable is neither long-run excluded (Juselius, 1995), nor weakly exogenous (Johansen, 1992).[6]

[5] As can be seen from their simulations, in the near cointegration case the true DGP is a first order vector moving average model that exhibits considerable size distortion with samples as large as $T = 400$ observations. This does not go away as the sample evolves.

[6] The empirical results are obtained using CATS 2.0 in RATS (see Dennis et al., 2005).

3.2. Data

For this paper, we use quarterly seasonally unadjusted data, where available, for the United States vis-à-vis Japan over the period 1980:1–2009:4. We choose the start of the sample period in order to control for structural change in the Japanese financial system because by the end of 1979, the interbank rates in Japan were deregulated, capital controls were removed, and the certificate of deposit market developed (McCornac, 1991). We use quarterly data as GDP data are not available on a monthly basis. The short-term interest rates are represented by the official discount rates,[7] whereas the long-term interest rates are represented by the 10-year government bond yields. Moreover, we use the consumer price index (CPI) to deflate the stock price indices represented by the S&P 500 in the United States and the Nikkei 225 in Japan.

While government spending is defined as government consumption in proportion to GDP, the productivity is defined as industrial production divided by the corresponding employment level. The real oil price is the West Texas Intermediate (WTI) Cushing crude oil spot price (in dollars per barrel) deflated by the US CPI. The exchange rate (denoted as dollars per unit of yen), interest rates, national income, industrial production, and price levels (CPI) are sourced from the IMF's International Financial Statistics (IFS). Nonetheless, money supply (M1), oil price, and stock prices are from Thomson DataStream.[8] Government spending and employment figures, on the other hand, are obtained from the OECD main economic indicators (MEI).

4. Empirical results

A prerequisite for conducting cointegration tests is to check the time series properties of the variables under investigation as to their order of integration. The null of non-stationarity is tested using augmented Dicky–Fuller (ADF) tests (Dickey and Fuller, 1981) and DF-GLS tests of Elliott et al. (1996). The results, as displayed in Tables 1 and 2, indicate that all the variables require first differencing to be stationary, hence they are integrated of order one ($I(1)$).[9] Cointegration is then tested using the Johansen (1995) procedure. The first subsection presents the analysis of the RID model; the second subsection analyses the HM model; and finally validation of the hybrid model is reported in the third subsection.

4.1. Long-run analysis of the RID model

For the data set $X_{RID,t}$, the SBIC, HQIC, and AIC indicate that the VAR is first order ($p = 1$). However, in order to remove any serial correlation and enhance the specification of the model, we require $p = 4$. The Lagrange multiplier (LM) test for the presence of serial correlation and ARCH along with the Jarque–Bera test of non-normality is reported in Table 3 and suggests that, at the 5% level, there is no evidence of misspecification for the model. On the basis of this specification, the estimated eigenvalues and trace statistics are reported in Table 4.

The trace test indicates that the null hypothesis of no cointegration is rejected, but since the subsequent test, reported in Table 4, for the null ($r = 1$) does not exceed the critical value, this supports the idea that there is a single cointegrating vector. It follows that this vector can be

[7] The official discount rate has been for a long time a major policy instrument for the Bank of Japan and other short-term interest rates such as call rate and bill discount rates have moved in line with the official rate (Ueda, 1996).

[8] With regard to the oil price, it is available in DataStream from 1982 onwards. So, the preceding observations are obtained from the Federal Reserve Bank of St. Louis. The last month snapshot in each quarter is considered, hence these observations are consistent with the DataStream ones that are end of period.

[9] The optimal lag length is chosen on the basis of the general-to-specific procedure suggested by Hall (1994) for the ADF tests and the modified Akaike information criterion (MAIC) suggested by Ng and Perron (2001) for the DF-GLS tests.

Table 2
DF-GLS unit root test results.

	Level		First difference	
	Constant	Constant and trend	Constant	Constant and trend
e	0.087 (3)	−2.007 (3)	−3.177 (3)[a]	−2.578 (4)
$m-m^*$	−1.697 (6)	−1.209 (6)	−2.499 (3)[b]	−2.563 (3)
$y-y^*$	−0.694 (3)	−1.064 (3)	−1.967 (2)[b]	−3.380 (2)[b]
i^s-i^{s*}	−1.671 (1)	−2.792 (1)	−2.427 (2)[b]	−4.136 (2)[a]
i^l-i^{l*}	−0.975 (0)	−1.974 (0)	−1.962 (3)[b]	−3.529 (3)[b]
$s-s^*$	0.008 (0)	−1.177 (0)	−2.486 (6)[b]	−2.967 (6)
$gs-gs^*$	0.347 (3)	−1.516 (3)	−1.952 (2)[b]	−3.123 (2)[b]
$prod^T-prod^{T*}$	0.532 (5)	−2.497 (5)	−2.991 (4)[a]	−3.062 (4)[b]
$roil$	−0.982 (2)	−1.094 (2)	−11.71 (0)[a]	−11.81 (0)[a]

Note: The 1% and 5% critical values for the DF-GLS test are respectively −2.584 and −1.943 (without trend) and −3.557 and −3.011 (with the trend); the proper lag length, allowing for a maximum of eight lags, is selected by the modified Akaike information criterion (MAIC), represented in parentheses. [a] and [b] indicate significance at the 1% and 5% levels, respectively.

generically identified in the long-run via a normalisation process (see Boswijk, 1996), and from the literature, this implies a long-run relation to explain the nominal exchange rate based on the RID model (with t-statistics in parentheses) as follows:

$$e_t = \underset{(1.664)}{2.250}(m_t-m_t^*) + \underset{(3.301)}{14.988}(y_t-y_t^*) + \underset{(1.014)}{0.090}(i_t^s-i_t^{*s}) - \underset{(7.890)}{1.065}\left(i_t^l-i_t^{*l}\right). \tag{12}$$

By inspection of the above results the estimated coefficient on the relative money supply has the sign expected by theory, even though it is large relative to the hypothesised magnitude of 1. Moreover, based on one-sided inference, we consider it significant at the 5% level. However, the coefficients on the rest of the monetary fundamentals have signs that are not consistent with the theory, although relative income and the long-term interest rate differential are highly significant (at the 1% level).

To provide further insights into this long-run relation, we conduct long-run exclusion (LE), weak exogeneity (WE), and stationarity tests by imposing restrictions on α and β (see Johansen, 1995); the latter tests are conducted to provide further evidence with regard to the stochastic properties of the series. Even though it is often felt that the normalisation is innocuous, the significance of the LE test is informative of the likely appropriateness of a normalisation. According to Boswijk (1996), empirical identification generally requires satisfaction of further rank conditions. However, Burke and Hunter (2005; chapter 5) argue that any coherent strategy for identification ought to preclude normalisation on variables that are either long-run excluded or weakly exogenous. Cointegration is a property of two or more non-stationary series and thus normalisation is also inappropriate on stationary variables.

The tests of LE, WE, and stationarity are asymptotically distributed chi-squared (Johansen, 1992) and in Table 5 we report our results on a variable by variable basis. The LE tests are conducted by imposing a zero restriction on the relevant elements of β. If a zero restriction on an element of β for a specific variable is not rejected, then the long-run relation cannot be normalised on this variable. The WE tests, by contrast, are carried out by imposing a zero restriction on elements of α in turn. If a zero restriction on an element of α for a particular variable is not rejected, then this variable can be considered weakly exogenous; it drives the system instead of adjusting to it. The stationarity tests are conducted, under the null hypothesis of stationarity, in the multivariate setting by fixing each element in turn in a single cointegrating vector to unity and the remaining elements to zero.

Table 3
Misspecification tests of the RID model.

	LM (8)	ARCH (8)	Normality	Skewness	Ex. Kurtosis
Panel A: Single equation tests					
e	1.005 [0.438]	0.188 [0.992]	4.414 [0.110]	0.442	3.063
$m-m^*$	1.518 [0.163]	0.718 [0.674]	2.676 [0.262]	0.054	3.480
$y-y^*$	1.619 [0.131]	0.079 [0.999]	2.628 [0.268]	0.194	3.430
i^s-i^{s*}	0.497 [0.854]	0.609 [0.768]	5.308 [0.070]	−0.566	3.286
i^l-i^{l*}	0.678 [0.708]	1.139 [0.343]	0.916 [0.632]	−0.027	3.246
Panel B: System tests					
	1.204 [0.088]		17.25 [0.068]		

Notes: LM (8) is a Lagrange multiplier test of serial correlation up to order 8; p-values are reported in square brackets [.].

Table 4
Johansen cointegration test results for the RID model. The system comprises of [e, $m-m^*, y-y^*, i^s-i^{s*}, i^l-i^{l*}$].

$(p-r)$	r	Eigenvalue	Trace test	95% critical value	p-value
5	$r=0$	0.398	100.940	69.611	0.001[a]
4	$r\le 1$	0.186	42.054	47.707	0.526
3	$r\le 2$	0.119	18.219	29.804	0.832
2	$r\le 3$	0.021	3.534	15.408	0.965
1	$r\le 4$	0.009	1.053	3.841	0.344

Notes: The lag length is selected using the Akaike information criterion (AIC), subject to correction for serial correlation by the inclusion of further lags. r denotes the number of cointegrating vectors.
[a] Indicates significance at the 1% level.

As is evident from Table 5, the LE tests indicate that, except for the relative income and long-term interest rate differential, all the other variables can be excluded from the cointegrating relation. Hence, any long-run model based on the exchange rate may be ill defined, as the related parameter cannot be distinguished from zero. In the subsequent panel, the proposition that the exchange rate and short-term interest rate differential are weakly exogenous also cannot be rejected. This implies that, at best, the long-run relation ought to be conditioned on the exchange rate instead of being normalised on. Hunter (1992) among others presents similar findings for the exchange rate.

In conclusion, despite the existence of a long-run relation among the variables of the RID model, such a relation cannot explain the behaviour of the exchange rate as this variable can be both excluded and viewed as weakly exogenous for the cointegrating vector. However, the tests of stationarity following from the restriction mentioned before on the VAR support the proposition that the series are all difference stationary (see Table 5), in line with the results of single unit root tests in Tables 1 and 2.

4.2. Long-run analysis of the HM model

The findings given above cast serious doubt on the conventional monetary model regarding the dollar–yen exchange rate. Therefore, we consider it of paramount interest to investigate the reasons for this failure. To this end, the VAR model is now based on the vector X_{HM} that represents the hybrid version. Since the price of oil is a global factor

Table 5
Long-run exclusion (LE), weak exogeneity (WE), and stationarity (S) tests for the RID model.

	e	$(m-m^*)$	$(y-y^*)$	(i^s-i^{s*})	(i^l-i^{l*})
Panel A: LE tests					
χ^2 (1)	2.364	1.894	7.241	0.412	27.797
p-value	0.124	0.169	0.000[a]	0.521	0.000[a]
Panel B: WE tests					
χ^2 (1)	1.579	18.722	4.751	0.280	16.274
p-value	0.209	0.000[a]	0.029[b]	0.597	0.000[a]
Panel C: S tests					
χ^2 (1)	47.410	56.959	55.866	46.315	27.492
p-value	0.000[a]	0.000[a]	0.000[a]	0.000[a]	0.000[a]

Notes: The critical values with one cointegrating vector are 6.64 and 3.84 at the 1% and 5% levels, respectively. [a] and [b] indicate significance at the 1% and 5% levels, respectively.

Table 6
Misspecification tests for the HM model.

	LM (8)	ARCH (8)	Normality	Skewness	Ex. Kurtosis
Panel A: Single equation tests					
e	0.764 [0.634]	0.641 [0.741]	4.105 [0.128]	0.348	3.239
$m-m^*$	1.432 [0.197]	1.265 [0.270]	14.88 [0.000][a]	−0.243	5.745
$y-y^*$	1.277 [0.266]	0.866 [0.547]	1.320 [0.516]	0.298	2.973
i^s-i^{s*}	0.901 [0.519]	0.840 [0.569]	2.305 [0.315]	−0.343	3.349
i^l-i^{l*}	1.060 [0.398]	1.465 [0.179]	5.706 [0.057]	−0.307	2.748
$s-s^*$	1.969 [0.060]	0.915 [0.507]	2.672 [0.262]	0.300	3.431
$gs-gs^*$	1.913 [0.069]	0.901 [0.518]	1.990 [0.369]	−0.055	3.544
$prod^T-prod^{T*}$	0.765 [0.634]	0.455 [0.884]	41.80 [0.000][a]	0.371	7.317
Panel B: System tests					
	1.272 [0.052]		63.80 [0.000][a]		

Notes: LM (8) is a Lagrange Multiplier test of serial correlation up to order 8; p-values are reported in square brackets [.].
[a] Indicates significance at the 1% level.

and all other factors are differentials between the United States and Japanese variables, we treat the real oil price as exogenous to the system.[10] Indeed, the test suggests a non-rejection of the null hypothesis of weak exogeneity with a p-value of 0.741. This finding is also consistent with the intuition of Amano and van Norden (1998) that oil prices in the decades preceding their study were governed by the major supply-side shocks resulting from political instability in the Middle East, and are thus external to the developed economies.

With regard to the VAR specification, the SBIC, HQIC and AIC suggest a lag length $p = 1$, while diagnostic tests imply that $p = 3$ is required to improve the specification. The reported diagnostics in Table 6 suggest that, at the 5% level, the model does not suffer from serial correlation using the LM test up to order 8, and the same applies for ARCH effects up to order 8. However, the multivariate normality test is rejected, where the sources of such failure seem to result from excess kurtosis in the money supply and productivity differentials. Since Gonzalo (1994) demonstrated a lack of sensitivity of the cointegrating rank to excess kurtosis, we conclude that these findings are robust.

Accordingly, Table 7 reports the trace test related to the HM model. It is evident that the null hypothesis of no cointegration is rejected, but evidence for more than one cointegrating vector cannot be rejected at the 5% level. Since the cointegrating rank does not change by the inclusion of the augmenting factors, this indicates that these factors follow stochastic trends common to the nominal exchange rate and its monetary fundamentals in the RID model. Long-run exclusion tests are likely to give more information regarding the nature of the contribution of the augmenting factors and also the variables on which the long-run relation may be normalised.

Hence, Table 8 reports the LE, WE, and stationarity tests of the variables included in the HM model. The stationarity tests imply that none of the variables in the cointegrating relation is stationary, consistent with the results reported in Tables 1 and 2. The LE tests, by contrast, indicate that the real oil price is the primary candidate for exclusion in the long-run relation, while the money supply and real stock price differentials could be excluded on a single-variable basis, although this would be rejected at the 15% level. However, at this stage, we do not exclude any variable based on a single-variable test. In the next subsection, we use these results to obtain a more parsimonious long-run relation.

Our key findings are that the long-run exclusion of the nominal exchange rate is rejected now, and that the exchange rate appears not to be weakly exogenous for the HM model (see panel B in Table 8). The change in WE status is a de facto indication of changes in long-run feedback and is of paramount interest (Juselius and Macdonald, 2004). Unlike the RID model, this finding indicates that the nominal exchange rate in the HM model adjusts to the long-run equilibrium. That is, it does not force the system when such a system accounts for the relative real stock prices, the productivity differential, relative government spending, and the real oil price. In addition to the real oil price on which the system is conditioned as stated earlier, the tests reported in Table 8 (panel B) also indicate that we cannot reject the findings that relative money supply, relative income, short-term interest rate differential, relative real stock prices, productivity differential, and relative government spending are weakly exogenous at the 5% level for the long-run relation, although the long-term interest rate differential is not.

Overall, the findings shown in Table 8 suggest that the long-run relation can be normalised on the nominal exchange rate, in line with the monetary model of the exchange rate, and primarily driven by the real and financial market shocks, corresponding to the results of Ahn and Kim (2010). Note that the findings on long-run weak exogeneity for both the short-term and long-term interest rate differentials do not vary according to model specification whether it is RID or HM and are consistent with the term structure of interest rates. This piece of information is important for the conduct of monetary policy because findings on the term structure are not supportive when the interest rate data are analysed alone. It follows that this relation is identified as a nominal exchange rate equation (with t-statistics in parentheses):

$$\begin{aligned} e_t = {} & \underset{(1.77)}{0.935}(m_t-m_t^*)-\underset{(3.61)}{5.524}(y_t-y_t^*)-\underset{(5.46)}{0.214}(i_t^s-i_t^{*s})+\underset{(4.83)}{0.262}\left(i_t^l-i_t^{*l}\right) \\ & -\underset{(2.52)}{0.477}(s_t-s_t^*)-\underset{(5.52)}{7.822}(prod^T{}_t-prod_t{}^{T*})-\underset{(8.34)}{12.42}(gs_t-gs_t{}^*) \\ & +\underset{(1.25)}{0.205}(roil_t). \end{aligned} \tag{13}$$

As shown from Eq. (13), the estimated coefficients on monetary fundamentals (relative money supply, relative income, and short-term and long-term interest rate differentials) are all significant and consistent with monetary theory. More specifically, the coefficient on the relative

Table 7
Johansen cointegration test results for the HM model. The system comprises of [e, $m-m^*$, $y-y^*$, i^s-i^{s*}, i^l-i^{l*}, $s-s^*$, $gs-gs^*$, $Prod^T-Prod^{T*}$, $roil$].

$(p-r)$	r	Eigenvalue	Trace test	95% critical value	p-value
8	$r=0$	0.443	230.054	204.989	0.002[a]
7	$r\le 1$	0.389	161.594	166.049	0.085
6	$r\le 2$	0.245	103.946	131.097	0.630
5	$r\le 3$	0.191	71.056	100.127	0.799
4	$r\le 4$	0.144	46.276	73.128	0.864
3	$r\le 5$	0.121	28.068	50.075	0.869
2	$r\le 6$	0.095	12.942	30.912	0.917
1	$r\le 7$	0.011	1.259	15.331	0.998

Notes: The lag length is selected using the Akaike information criterion (AIC), subject to correction for serial correlation by the inclusion of further lags. r denotes the number of cointegrating vectors.
[a] Indicates significance at the 1% level.

[10] Johansen and Juselius (1992) assume that the real oil price is strictly exogenous. Hunter (1992) shows that this corresponds in the long-run to the oil price being weakly exogenous and long-run excluded, but these restrictions were found to be rejected.

Table 8
Long-run exclusion (LE), weak exogeneity (WE), and stationarity (S) tests for the HM model.

	e	$(m-m^*)$	$(y-y^*)$	(i^s-i^{s*})	(i^l-i^{l*})	$(s-s^*)$	$(prod^T-prod^{T*})$	$(gs-gs^*)$	$roil$
Panel A: LE tests									
χ^2 (1)	5.641	2.058	4.556	9.756	7.942	2.561	6.057	9.163	0.845
p-value	0.018[b]	0.150	0.033[b]	0.002[a]	0.005[a]	0.110	0.014[b]	0.002[a]	0.358
Panel B: WE tests									
χ^2 (1)	3.978	0.192	0.225	0.006	4.220	2.461	1.627	3.261	
p-value	0.046[b]	0.662	0.636	0.939	0.040[b]	0.117	0.202	0.071	
Panel C: S tests									
χ^2 (1)	38.038	38.254	37.133	39.459	23.569	30.839	29.992	33.191	
p-value	0.000[a]	0.000[a]	0.000[a]	0.000[a]	0.000[a]	0.000[a]	0.000[a]	0.000[a]	

Notes: The critical values with one cointegrating vector are 6.64 and 3.84 at the 1% and 5% levels, respectively. [a] and [b] indicate significance at the 1% and 5% levels, respectively.

money supply is not materially different from 1, and is significant based on a one-sided test at the 5% level. All the other monetary variable coefficients (relative income and short-term and long-term interest rate differentials) have their hypothesised signs and are significant at the 1% level. Furthermore, as hypothesised by Frankel (1979), the parameter on the long-term interest rate differential is greater than that on the short-term interest rate differential in absolute value.

Except for the real oil price, all the factors that have been used to augment the monetary model have significant parameters. This implies that the real oil price can be excluded from the long-run relation and, as with Johansen and Juselius (1992), treated as strictly exogenous. Consistent with Friedman (1988) and Caruso (2001), the coefficient on the relative real stock price is negative implying that the wealth effect dominates the substitution effect in the underlying money demand functions for the United States and Japan. The coefficients on the productivity differential across the industrial sectors and relative government spending are negative and significant. This suggests that higher domestic productivity or government spending compared to their foreign counterpart results in an exchange rate appreciation. The fact that the real oil price can be excluded from the long-run part of the VAR system suggests that it affects the long-run only indirectly by enhancing the econometric performance of the model.

4.3. Hybrid model validation

The above results strongly indicate that the HM model dominates the RID model on theoretical and econometric grounds in explaining the dollar–yen exchange rate in the long-run. However, to check the robustness of our results, we conduct two further analyses. First, we use the results on LE and WE to obtain a more specific and robust formulation of the long-run relation based on the HM model. A similar approach has also been used by MacDonald and Nagayasu (1998), though they examine a simplified version of the RID model of the yen–dollar exchange rate excluding money demand functions.

Having determined that $r=1$, it follows that the structure of α and β, subject to $roil_t$ being weakly exogenous ($\alpha_8=0$), takes the following form:

$$\begin{array}{lccccccccc} & e & \nabla m & \nabla y & \nabla i^s & \nabla i^l & \nabla s & \nabla prod^T & \nabla gs & roil \\ \alpha' = [& \alpha_0 & \alpha_1 & \alpha_2 & \alpha_3 & \alpha_4 & \alpha_5 & \alpha_6 & \alpha_7 & 0], \end{array}$$

where ∇ represents the differential between the variables for the US and Japan. The cointegrating vector is normalised on the exchange rate by restricting ($\beta_0=-1$) and from the long-run exclusion tests we impose ($\beta_8=0$) for the real oil price:

$$\begin{array}{lccccccccc} & e & \nabla m & \nabla y & \nabla i^s & \nabla i^l & \nabla s & \nabla prod^T & \nabla gs & roil \\ \beta' = [& -1 & \beta_1 & \beta_2 & \beta_3 & \beta_4 & \beta_5 & \beta_6 & \beta_7 & 0]. \end{array}$$

Next, we sequentially impose zero restrictions on the loading factors, α, of the standard monetary fundamentals related to relative money supply, relative income, and short-term interest rate differential. These weak exogeneity restrictions are empirically plausible given the size of the adjustment coefficients and also consistent with monetary theory. The tests, as displayed in Table 9, indicate that the imposed restrictions cannot be rejected. Moreover, the constrained final long-run relation normalised on the exchange rate suggests the significance of monetary variables with their hypothesised signs, as found in the previous subsection. The trace test also implies that there is still a single cointegrating vector among the variables (these results are unreported). Overall, this demonstrates the robustness of our results in terms of the long-run formulation and direct impact of the augmenting factors on the long-run exchange rate monetary model.[11]

Then, we subject our proposed HM model to an array of forward and backward recursive stability tests proposed by Hansen and Johansen (1999) to gain further insights into its adequacy as a long-run exchange rate model. The results reported here relate to the behaviour of the max tests of β and are displayed in Fig. 2. The forward and backward tests appear respectively in the Figure's left and right panels, with the corresponding 5% critical value represented by the solid line. Note that in providing these stability tests, the short-run effects (X(t)) compared to those of the long-run R1(t) are concentrated out. In a broad sense, the model shows a reasonable degree of stability of the parameters in the cointegrating vector. Hence, the model seems to be adequate and does not exhibit structural breaks in relation to the long-run for the period under observation.

5. Conclusions

In this paper, we re-examine the dollar–yen exchange rate using two versions of the monetary model. The first is the conventional real interest differential (RID) model of Frankel (1979) and the second is a hybrid monetary (HM) model, proposed herein, that incorporates on the one hand real stock prices to capture the stability of money demand and on the other, the productivity differential, relative government spending, and the real oil price to explain the persistence in the real exchange rate. Both models are estimated using the Johansen cointegration methodology and quarterly data from 1980 to 2009, a period characterised by high international capital mobility, as well as periodic volatility in the dollar–yen exchange rate.

Although a single cointegrating vector exists for both models, the long-run exclusion and weak exogeneity tests inform us that the HM version gives an appropriate long-run explanation of the monetary

[11] To confirm that the failure of the monetary model is due to the breakdown of its underlying building blocks, we have also considered the performance of the standard flexible-price monetary model (Bilson, 1978) against its hybrid version. The results showed that there is no cointegration among the variables of the standard flexible-price monetary model (nominal exchange rate, relative money supply, relative income, and short-term interest rate differential). By contrast, the hybrid version that accounts for money demand instability and real exchange rate persistence, gives further support to the monetary model because it gives rise to a single long-run relation among the variables and coefficients consistent with the theory when the relation is normalised on the exchange rate. These results are available upon request from the authors.

Table 9
Joint tests of weak exogeneity and long-run exclusion conditional on $r = 1$ in the HM model.

Tests under the null:	Statistics [p-value]
(1) $\beta_8 = 0$	χ^2 (1) = 0.845 [0.358]
(2) $\beta_8 = 0, \alpha_3 = 0$	χ^2 (2) = 0.856 [0.652]
(3) $\beta_8 = 0, \alpha_3 = 0, \alpha_1 = 0$	χ^2 (3) = 1.690 [0.639]
(4) $\beta_8 = 0, \alpha_3 = 0, \alpha_1 = 0, \alpha_2 = 0$	χ^2 (4) = 2.010 [0.734]

The implied long-run relation by test (4):

	$(m - m^*)$	$(y - y^*)$	$(i^s - i^{s*})$	$(i^l - i^{l*})$	$(s - s^*)$	$(prod^T - prod^{T*})$	$(gs - gs^*)$
Coefficients	0.740	−4.028	−0.169	0.172	−0.557	−6.748	−11.231
t-statistics	−1.743[b]	3.363[a]	5.758[a]	4.529[a]	3.669[a]	5.743[a]	9.095[a]

Notes: [a] and [b] indicate significance at the 1% and 5% levels, respectively; p-values are in square brackets [.]. The coefficient on the relative money supply is significant at the 5% level based on a one-sided test.

model of the dollar–yen exchange rate. The enhanced performance of the HM model derives from the following considerations to the conventional monetary model. First, the stability of money demand relations is taken into account by the inclusion of key variables that impact on transactions (Friedman, 1988). A key feature of globalised financial markets is a highly active market in cross-border investments, mergers and acquisitions, and cross-listed stocks. In particular, the futures contract on the Nikkei is listed as an asset in the US stock market. Second, the persistence of the real exchange rate, which reflects primarily the impact of the non-traded goods, is taken into consideration by accounting for productivity and government expenditure differences. In essence, these differences may be due to the relatively insular nature of Japanese society limiting the effectiveness of arbitrage. The literature also suggests that the real oil price affects such persistence, but the empirical findings herein show an indirect impact of such a price via the dynamic specification of the VAR model.

Contrary to the conventional monetary model, the results also suggest that the dollar–yen exchange rate in the hybrid model is driven by money, income, and short-term interest rate differentials, but not the reverse. This implies a substantial role for real economic and financial market variables in a well-formulated monetary model for the determination of the long-run exchange rate.

Acknowledgements

We would like to thank Stephen Hall (the editor) and two anonymous referees of this journal for their very useful comments and suggestions. We are also grateful to Ray Barrell, Sugata Ghosh, Guglielmo Maria Caporale, Panagiotis Konstantinou, Matthew Greenwood Nimmo, the participants at the EEFS12 conference at Koç University (Istanbul, Turkey), including George Chortareas, Sushanta Mallick and Keith Pilbeam, and the participants at the MMF Conference at the University of Birmingham in May 2012 for their very helpful comments and suggestions on an earlier version of this paper, which was previously circulated with the title “The Monetary Model of the US Dollar–Japanese Yen Exchange Rate: An Empirical Investigation”. All remaining errors are our own responsibility.

Appendix A. Derivation of the HM model

Following Friedman (1988), the money demand equations in Frankel's (1979) RID model, Eqs. (3) and (4) above, are modified as follows:

$$m - p = a_1 y - a_2 i + a_3 s, \tag{1a}$$

$$m^* - p^* = a_1 y^* - a_2 i^* + a_3 s^*, \tag{2a}$$

where m is the money supply, p the price level, i the interest rate, y real income, and s the real stock price. All variables except interest rates are in logs, and the asterisk denotes the foreign economy. Besides the income elasticity and the interest rate semi-elasticity of money demand (as discussed above), for simplicity the real stock price elasticity of money demand, a_3, is also assumed to be identical across the domestic and foreign countries.

If, in addition to the assumption of UIP, PPP also holds along with $\left(\bar{i} - \bar{i}^*\right) = \left(\Delta p^e - \Delta p^{*e}\right)$ in the long-run, then Eq. (7) above is modified (where a bar denotes an equilibrium value) as follows:

$$\begin{aligned} e = {} & \left(\overline{m} - \overline{m}^*\right) - a_1\left(\overline{y} - \overline{y}^*\right) \\ & + a_2\left(\Delta p^e - \Delta p^{*e}\right) - \frac{1}{\theta}\left[\left(i - \Delta p^e\right) - \left(i^* - \Delta p^{*e}\right)\right] + a_3\left(\overline{s} - \overline{s}^*\right). \end{aligned} \tag{3a}$$

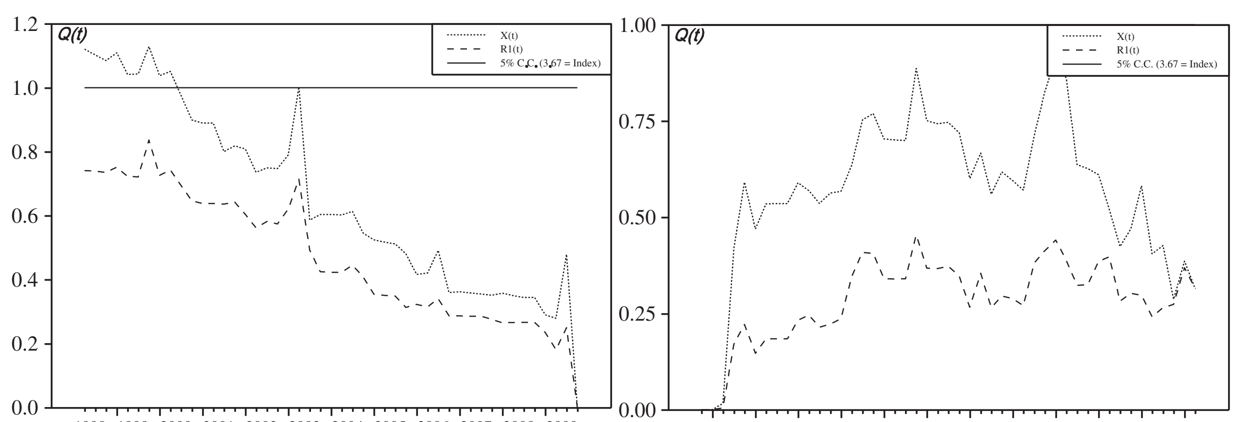

Fig. 2. Forward (left panel) and backward (right panel) recursively calculated test for the constancy of β in the hybrid monetary model (1.0 corresponds to the 5% critical value).

However, as shown from Fig. 1 and discussed earlier, the real exchange rate is evidently persistent and as a result PPP is not tenable. To account for the impact of real economic shocks on this type of persistence as documented in the literature, then following Clements and Frenkel (1980) we decompose the aggregate price levels into the prices of traded $\overline{p}^T$ and non-traded $\overline{p}^{NT}$ goods as follows:

$$\overline{p} = (1-a)\overline{p}^T + a\overline{p}^{NT} = \overline{p}^T + a\left(\overline{p}^{NT} - \overline{p}^T\right), \tag{4a}$$

$$\overline{p}^* = (1-a)\overline{p}^{*T} + a\overline{p}^{*NT} = \overline{p}^{*T} + a\left(\overline{p}^{*NT} - \overline{p}^{*T}\right), \tag{5a}$$

where $a(1-a)$ denotes the proportion of non-traded (traded) goods in the economy. Meanwhile, the real exchange rate $\overline{q}$ is the nominal exchange rate adjusted for domestic and foreign price levels:

$$\overline{q} = \overline{e} - \overline{p} + \overline{p}^*. \tag{6a}$$

Substituting the aggregate price levels in Eq. (6a) by those in Eqs. (4a) and (5a), then the real exchange rate is:

$$\overline{q} = \left(\overline{e} - \overline{p}^T + \overline{p}^{*T}\right) - a\left[\left(\overline{p}^{NT} - \overline{p}^T\right) - \left(\overline{p}^{*NT} - \overline{p}^{*T}\right)\right]. \tag{7a}$$

If PPP applies primarily to traded goods, then $\left(\overline{e} - \overline{p}^T + \overline{p}^{*T}\right)$ in Eq. (7a) should be zero (see Schnabl, 2001, for the validity of PPP for Japan using traded goods) and the real exchange rate expressed in terms of both traded and non-traded goods is:

$$\overline{q} = -a\left[\left(\overline{p}^{NT} - \overline{p}^T\right) - \left(\overline{p}^{*NT} - \overline{p}^{*T}\right)\right]. \tag{8a}$$

In a competitive world, the price in each sector should reflect the unit labour costs in the sector, and as Strauss (1999) pointed out this will clarify the relative price movements of non-traded goods, so that:

$$\overline{p}^T = \overline{w} - \overline{prod}^T, \overline{p}^{NT} = \overline{w} - \overline{prod}^{NT}, \overline{p}^{*T} = \overline{w}^* - \overline{prod}^{*T}, \overline{p}^{*NT} = \overline{w}^* - \overline{prod}^{*NT}, \tag{9a}$$

where $\overline{w}$ is the wage rate equated across both the traded and non-traded sectors due to internal labour mobility, while $\overline{prod}^T$ ($\overline{prod}^{NT}$) indicate the productivity in the traded (non-traded) sectors. This implies the following:

$$\overline{p}^{NT} - \overline{p}^T = \overline{prod}^T - \overline{prod}^{NT}, \overline{p}^{*NT} - \overline{p}^{*T} = \overline{prod}^{*T} - \overline{prod}^{*NT}. \tag{10a}$$

Substituting the expressions in Eq. (10a) into Eq. (8a) results in the following real exchange rate relation:

$$\overline{q} = -a\left[\left(\overline{prod}^T - \overline{prod}^{NT}\right) - \left(\overline{prod}^{*T} - \overline{prod}^{*NT}\right)\right]. \tag{11a}$$

To further capture the impact of the demand side shocks proxied by government spending (see Chinn, 2000) and the terms of trade shocks represented by the real oil price (see Amano and van Norden, 1998), we extend Eq. (11a) as follows:

$$\overline{q} = -a\left[\left(\overline{prod}^T - \overline{prod}^{NT}\right) - \left(\overline{prod}^{*T} - \overline{prod}^{*NT}\right)\right] + \lambda\left(\overline{gs} - \overline{gs}^*\right) + \delta\overline{roil}, \tag{12a}$$

where $\overline{gs}$ ($\overline{gs}^*$) denotes domestic (foreign) government consumption as a percentage of GDP and $\overline{roil}$ is the real oil price (where the US CPI is used to deflate the oil price). Chinn (1997) explains that quarterly data for the non-traded sector is limited, and this leads to the assumption that $\overline{prod}^{NT} = \overline{prod}^{*NT}$. As a result, Eq. (12a) can now be expressed as follows:

$$\overline{q} = -a\left(\overline{prod}^T - \overline{prod}^{*T}\right) + \lambda\left(\overline{gs} - \overline{gs}^*\right) + \delta\overline{roil}. \tag{13a}$$

Using the expression (13a) in Eq. (3a), we obtain the HM model:

$$e = \left(\overline{m} - \overline{m}^*\right) - a_1\left(\overline{y} - \overline{y}^*\right) + a_2\left(\Delta p^e - \Delta p^{*e}\right) - \frac{1}{\theta}\left[\left(i - \Delta p^e\right) - \left(i^* - \Delta p^{*e}\right)\right] + a_3\left(\overline{s} - \overline{s}^*\right) - a\left(\overline{prod}^T - \overline{prod}^{*T}\right) + \lambda\left(\overline{gs} - \overline{gs}^*\right) + \delta\overline{roil}. \tag{14a}$$

Eq. (14a) is estimated on the basis that the short-term interest rates represent real interest rates and long-term interest rates proxy long-run expected inflation rates, as discussed above. That is written as follows:

$$e_t = \beta_1\left(m_t - m_t^*\right) + \beta_2\left(y_t - y_t^*\right) + \beta_3\left(i_t^s - i_t^{s*}\right) + \beta_4\left(i_t^l - i_t^{l*}\right) + \beta_5\left(s_t - s_t^*\right) + \beta_6\left(prod_t^T - prod_t^{T*}\right) + \beta_7\left(gs_t - gs_t^*\right) + \beta_8 roil_t + v_t,$$

where $\beta_1, \beta_2, \beta_3, \beta_4, \beta_5, \beta_6, \beta_7$, and β_8 represent the unrestricted parameters on the respective variables.

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