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Determinants of real exchange rates: An empirical investigation

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Contents

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Abstract

The large and persistent deviations of nominal exchange rates from their purchasing power parities comprise a key stylized fact in international economics. This paper sheds light on these persistent deviations by combining two disparate strands of empirical work. The first strand focuses on real economic shocks such as sectoral technology shocks suggested by the celebrated Balassa-Samuelson model, whereas the second strand emphasizes monetary shocks which create persistent effects on both the real interest rate and the real exchange rate. We also hypothesize a third factor which may affect real exchange rates – shocks to the global financial system, which we proxy by the real price of gold.

Although each factor in isolation has limited explanatory power, we find that these three factors in conjunction can successfully explain the medium to long run movements in 14 bilateral U.S. dollar real exchange rates from 1970 to 2006. The three factors are sectoral total factor productivity differentials, real interest rate differentials, and the real price of gold, representing real shocks, monetary shocks, and shocks to the global financial system, respectively. We document evidence suggesting that bilateral U.S. dollar real exchange rates are cointegrated with these three factors.

Keywords: purchasing power parity, Balassa-Samuelson model, cointegration

JEL codes: F31, F41

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1 Introduction

The real exchange rate, defined as the price level of the home country relative to the foreign country and adjusted by the nominal exchange rate, is a key variable in international macroeconomics. One of the most widely documented and robust stylized fact in international finance is that changes in real exchange rates tend to be very persistent, and real exchange rates follow a time series process which is well-approximated by a unit-root non-stationary process.[1] Put another way, the deviations between nominal exchange rates and their Purchasing Power Parities (PPPs) tend to be large and very persistent.

The most celebrated explanation for the near permanent changes in real exchange rates is the so-called Balassa-Samuelson effect (see Balassa (1964) and Samuelson (1964)). The key insight behind the Balassa-Samuelson effect is that productivities evolve differently in the traded- and non-traded-goods sectors, as a consequence of which the relative price of non-traded goods changes over time, thereby altering the equilibrium real exchange rate in the long run.[2]

By constructing sectoral total factor productivities (*TFPs*) for a panel of 15 OECD countries as well as China, this paper investigates whether differential trends in sectoral *TFPs* in the home and foreign countries can explain the stochastic trend in bilateral real exchanges.[3] In addition to sectoral TFP differentials, we also consider two other variables which could have an important influence in understanding the evolution of real exchange rates – real interest rate differentials and the price of gold.

There is a large literature examining the relationship between real exchange rates and real interest rate differentials. For instance, Messe and Rogoff (1988) and Edison and Pauls (1993), examine whether real exchange rates are cointegrated with real interest rate differentials, but are unable to find conclusive evidence of such a relationship. Baxter (1994) examines the link between real exchange rates and real interest rate differentials at

[1] For instance, Baillie and McMahon (1989), Corbae and Ouliaris (1988), Engel (2000), Mark (1990), Meese and Rogoff (1988), O'Connell (1998), and several other researchers document evidence suggesting that real exchange rates are unit-root nonstationary, thereby implying that PPP fails to hold even in the long-run. While several studies (e.g. Frankel and Rose 1995, Lothian and Taylor 1996) have also found evidence of mean reversion in real exchange rates, the slow rate of convergence to PPP suggests that persistent real disturbances could plausibly be playing an important role. See the surveys by Froot and Rogoff (1995) and Rogoff (1996) for further details.

[2] Obstfeld and Rogoff (1996, Chapter 4) provides an excellent overview of the theory and evidence on the Balassa-Samuelson model.

[3] The “home” country is typically an OECD country, whereas the “foreign’ country is always the U.S., so that these are bilateral U.S. real exchange rates.

various frequencies, and finds that the relationship is strongest at the medium- and low-frequencies.

Following a somewhat different line of enquiry, Eichenbaum and Evans (1995) investigate the effects of shocks to U.S. monetary policy on exchange rates. They document robust evidence showing that a contractionary shock to U.S. monetary policy leads to sharp and persistent increases in U.S. interest rates, and the same shocks also lead to persistent appreciations in U.S. nominal and real exchange rates. Indeed, they show that following a contractionary shock, the dollar continues to appreciate for a substantial period of time.[4] Bacchetta and van Wincoop (2010) explain this phenomenon of "delayed overshooting" by formalizing Froot and Thaler's (1990) intuition of infrequent portfolio decisions on the part of investors.

It is interesting to note that the empirical work on the Balassa-Samuelson effect and the relationship between interest rate differentials and real exchange rates has largely proceeded independently. Thus, one of the reasons why earlier work could not detect a cointegrating relationship between real exchange rates and real interest rate differentials could be due to omitted variable bias due to the highly persistent sectoral productivity differentials being ignored in the analysis. Likewise, the empirical literature on the Balassa-Samuelson model has generally ignored the persistent effects of monetary shocks on real exchange rates.[5] A combination of the two effects is expected to shed greater light on the evolution of the real exchange rate.

The third factor we consider in our analysis of U.S. bilateral real exchange rates is the real price of gold. While sectoral TFP differentials capture the effects of real shocks to real exchange rates, and real interest rates capture the effects of monetary shocks, a third possible source of shocks would be those to the international financial system. To the extent that the U.S. dollar is the premier reserve currency of the global economy, episodes of severe stress to the global financial system, such as those witnessed in the crisis of 2008−09, would inevitably have an impact on the value of the dollar as investors' confidence in fiat-money is shaken. To the extent that gold is regarded as a hedge against the

[4] Engel (2012) also discusses the evidence for the stylized fact that high real interest rates are accompanied by strong and appreciating real exchange rates.

[5] Recent empirical work on the Balassa-Samuelson model, such as Thomas and King (2008) and Kakkar and Yan (2012) has focused on relatively rapidly growing Asia-Pacific economies, whereas earlier work, such as De Gregorio, Giovannini, and Wolf (1994), Chinn and Johnston (1996), Canzonerri, Cumby and Diba (1999), Kakkar and Ogaki (1999) and Kakkar (2003) studied largely the OECD economies. While some of these papers consider variables other than TFP differentials, none of them consider real interest rate differentials.

collapse of fiat money, the real price of gold can serve to some extent as a proxy for shocks to the global financial system. For example, from the end of 2008 to the end of 2009, as the financial crisis originated from and deepened in the U.S., the U.S. dollar lost nearly 17% of its value against a basket of currencies, while the real price of gold increased about 25%. To the best of our knowledge, the connection between real exchange rates and real gold prices has not been explored in the extant empirical literature.

Our main findings are as follows. In contrast to the earlier literature on the Balassa-Samuelson hypothesis, we find that for 10 out of 14 countries in our panel, real exchange rates are cointegrated with sectoral TFPs. Including real interest rate differentials and the real world price of gold makes the cointegration results much stronger, and noticeably increases the *p*-values for the null hypothesis of cointegration. Moreover, the coefficients of these 3 factors are generally consistent with economic theory as well as statistically significant.

The rest of the paper is organized as follows. The next section describes the empirical specification motivated by the Balassa-Samuelson model, and its augmented versions, which include real interest rate differentials and the real price of gold. Section III describes the data used in this paper. Section IV presents the empirical results and Section V concludes.

2 Model specification

Consider a world economy with two countries − country *H* is the home country and country *F* is the foreign country. Each country is an open economy that is divided into traded- and nontraded-goods sectors: good *T* is tradable and good *N* is nontradable. Both goods are produced via Cobb-Douglas production functions. The production side of the economy can be described as follows:

$$Y_T^i = A_T^i (L_T^i)^{\alpha_T^i} (K_T^i)^{1-\alpha_T^i} \quad , i = H, F \tag{1}$$

$$Y_N^i = A_N^i (L_N^i)^{\alpha_N^i} (K_N^i)^{1-\alpha_N^i} , \; i = H, F \tag{2}$$

Here *Y* denotes output; *L* and *K* denote labor and capital, respectively; *A* denotes *TFP* and α denotes the share of labor in production.

Under the standard assumptions of the Balassa-Samuelson model[6] one obtains the following expression for the relative price of nontraded-goods in terms of the exogenous sectoral *TFP*s:

$$ln(Q^i) = \lambda^i + (\alpha_N^i / \alpha_T^i) ln(A_T^i) - ln(A_N^i) = \lambda^i + d^i \qquad (3)$$

Here λ^i is a constant that depends on the sectoral labor shares, and $d^i = (\alpha_N^i / \alpha_T^i) ln(A_T^i) - ln(A_N^i)$ is country *i*'s labor-share-adjusted sectoral TFP differential. Equation (3) shows that the relative price of nontraded-goods in each country is determined by the sectoral *TFP* differentials.

We assume that the price level of each country P^i ($i = H, F$) can be approximated by a geometric average of the prices of nontradable and tradable goods up to a stationary measurement error:

$$P^i = c^i (P_N^i)^{\beta^i} (P_T^i)^{(1-\beta^i)}. \qquad (4)$$

Here β^i is the share of nontradables in the overall price level of country *i* and c^i is a stationary measurement error that reflects factors which cause the general price level to deviate from the geometric average of the price of non-tradable and tradable goods. Let *E* denote the nominal exchange rate between the home and foreign countries. *E* units of the home country's currency buy 1 unit of the foreign currency. The real exchange rate between the home and foreign countries, E^r, is the ratio of the home price level to the foreign price level adjusted by the nominal exchange rate:

$$E^r = \frac{P^H}{EP^F} \qquad (5)$$

If the law of one price holds for tradable goods in both countries, then there exists a relationship between the real exchange rate and the relative price of nontradables. Given the

[6] See, for example, Chapter 4 of Obstfeld and Rogoff for an excellent exposition of the Balassa-Samuelson model, or Kakkar and Yan (2012) for a modern time-series interpretation.

presence of transportation costs and other frictions in the international trade process, the law of one price may not hold in the short run. This paper makes the more plausible assumption that the law of one price holds for tradable goods in the long run, so that the real exchange rate for tradable goods is stationary. We can express this assumption mathematically via the following equation:

$$\ln(P_T^H) = \ln(E) + \ln(P_T^F) + u, \quad (6)$$

where u is a stationary random variable. The stationarity of u implies that deviations from PPP for tradables do not persist forever and are transient. Combining equations (4) to (6), we can express the bilateral real exchange as a linear combination of the relative price of nontradables in the home and foreign countries:

$$\ln(E^r) = \mu + \beta^H \ln(Q^H) - \beta^F \ln(Q^F) + \epsilon, \quad (7)$$

where $\mu = \{E(\ln(c^H) - E(\ln(c^F)\}$ is a constant and $\epsilon = u + \{\ln(c^H) - E(\ln(c^F)\} - \{\ln(c^F) - E(\ln(c^F)\}$ is a zero-mean stationary random variable. Finally, we combine equations (3) and (7) to express the real exchange rate in terms of home and sectoral productivity differentials:

$$\ln(E^r) = \theta + \beta^H d^H - \beta^F d^F + \varepsilon, \quad (8)$$

where $\theta = (\mu + \lambda^H - \lambda^F)$ is a constant.

Equation (8) encapsulates the key prediction of the Balassa-Samuelson model, which is that the bilateral real exchange rate is determined by sectoral *TFP* differentials in the home and foreign countries. Since bilateral real exchange rates are well approximated by a near-unit root process, the model implies that different stochastic trends in the home and foreign sectoral *TFP* differentials lead to persistence in the real exchange rates. Hence, model predicts that real exchange rates should be cointegrated with the home and foreign sectoral *TFP* differentials. Equation (8) forms our benchmark cointegrating regression.

As mentioned in the Introduction, the empirical literature has documented significant and persistent effects of real interest rate differentials on real exchange rates, and we also hypothesize that the real price of gold may serve as a proxy for shocks to the global financial system. Hence in addition to the benchmark regression, we also estimate the following augmented cointegrating regression:

$$\ln(E^r) = \theta + \beta^H d^H - \beta^F d^F + \beta_r r_{diff} + \beta_g \ln\left(rp_{gold}\right) + \varepsilon', \quad (9)$$

where r_{diff} represents the real interest rate differential between the home and foreign countries, and rp_{gold} represents the real price of gold.

3 Data

We collected industry level data on output, labor compensation, working hours, investment and capital from 1970−2006 for 15 OECD countries as well as China. The sectoral data for the OECD countries were primarily from the STAN database (Structural Analysis Database) of OECD. The following industries were classified as tradable: agriculture, hunting and fishing; manufacturing; mining and quarrying; wholesale, retail trade, restaurants and hotel; and transport, storage and communication. The nontradable sector comprised: electricity, gas and water; construction; finance, insurance, real estate and business services; and community, social and personal services.[7] Sectoral *TFP* was constructed as a Solow-residual ((Solow (1957)) from constant-price domestic currency series of output, capital, labor shares and hours worked in that sector. Nominal values were converted to the real values using the industry level implicit GDP deflators.

Corresponding data from China were collected from the CEIC database. The classification of tradable goods in China included the primary sector, "industry" in the secondary sector as well as the "wholesale, retail and catering" and "transportation, cost and telecom" in the tertiary sector. Nontradable goods for China included "construction" in secondary sector and other industries in the tertiary sector.

Where sectoral capital stocks were not directly available from the databases, they were estimated using the perpetual inventory method.[8] Where total working hours were not directly available from the databases, they were estimated by multiplying the "total em-

[7] Our classification is the same as that commonly used for OECD countries in the empirical literature (see, for example, by De Grogorio, Giovannini, and Wolf (1994)) and Stockman and Tesar (1995)

[8] The initial capital stock was taken to be five times the real gross fixed capital formation in the base year. The fixed capital stock was assumed to have an annual depreciation rate of 5%. New real gross fixed capital formation was then added to the capital stock brought over from last period.

ployment" with the "average hours worked per person engaged" which is reported in the GGDC (Groningen Growth and Development Centre) database.

National level implicit GDP deflators for various countries were obtained primarily from the World Development Indicators database of the World Bank and the OECD iLibrary, supplemented by data from various national statistical agencies. Data for short-term and long-term interest rates were collected from the International Financial Statistics (IFS), OECD iLibrary, CEIC, Datastream and the webpages of the central banks of various countries (details are provided in the appendix). Exchange rate data was primarily collected from the IFS and supplemented by the CEIC database and Eurostat. Gold price data was taken from Bloomberg.

4 Empirical results

We begin our analysis by examining the evidence for stochastic trends in the dependent and explanatory variables. In addition to the Phillips-Perron test (Phillips and Perron, 1988) we also use the $J(1, 5)$ test of Park (1990) to test the null hypothesis of unit-root stationarity against the alternative of trend stationarity.

Panel A of Table 1 reports the evidence for stochastic trends in real exchange rates, sectoral *TFP* differentials and real interest rate differentials. For bilateral real exchange rates against the U.S. dollar, the J(1, 5) test does not reject the null hypothesis of a unit root for any country except the United Kingdom, for which it is significant at the 5% level. The Phillips-Perron test without trend is significant at the 10% level for Finland and the Netherlands, but not for any other country. The Phillips-Perron test with trend is not significant for any of the real exchange rates even at the 10% level. These results are consistent with a large body of evidence documenting the near unit-root behavior of real exchange rates.

For sectoral *TFP*-differentials, the $J(1, 5)$ test rejects the null of a unit for South Korea and Sweden at the 1% level. However, the Phillips-Perron test does not reject the null hypothesis of a unit root for any of the countries, regardless of whether a time trend is included under the alternative.

For the real interest rate differential between the home country and the U.S., the $J(1, 5)$ test rejects the null hypothesis of a unit root for Belgium and France at the 5% level. The Phillips-Perron test without trend, which is more relevant for this case, does not reject the null hypothesis of a unit root for any of the countries. The Phillips-Perron test with trend is significant for Belgium, Denmark and France, at the 10%, 5% and 1% significance levels, respectively.

Panel B of Table 1 contains the unit root tests for the real price of gold. The Phillips-Perron test without trend is significant at the 10% level, but both the J(1, 5) test and the Phillips-Perron test with trend do not reject the null hypothesis of a unit root.

To summarize, the evidence presented in Table 1 is largely consistent with the idea that the time series processes followed by real exchange rates, the composite productivity differential, the real interest rate differential and the real price of gold are well-approximated by processes which are unit-root nonstationary.

Having verified that the pre-condition for testing cointegration is satisfied in our data, we now test for cointegration between real exchange rates and our set of explanatory variables. We employ the Canonical Cointegrating Regressions (*CCR*) procedure proposed by Park (1992) to test the null hypothesis of cointegration predicted by the Balassa-Samuelson model. The CCR estimators are asymptotically efficient and have asymptotic distributions that can be essentially considered as normal distributions, so that their standard errors can be interpreted in the usual way.[9]

The CCR procedure is applied to the regression

$$y(t) = \theta_c + \sum_{i=1}^{p} \eta_i t^i + \sum_{i=p+1}^{s} \eta_i t^i + \gamma X(t) + \varepsilon_c(t), \tag{10}$$

where $y = ln(E^r)$, and X is the vector of explanatory variables. In order to test for stochastic and deterministic cointegration, superfluous time polynomials up to order s are included in this regression. When the estimated cointegrating vector has been incorporated in the canonical cointegrating regression, if all the coefficients of the added time polynomials are zero, the deterministic cointegration restriction is satisfied. If all coefficients of the added time polynomials are zero, with the exception of the first-order time polynomial, stochastic

[9] The CCR estimators also have the advantage that they do not require the assumption of a Gaussian VAR structure. Monte Carlo experiments in Park and Ogaki (1991) show that the CCR estimators have better small sample properties than Johansen's (1988; 1991) estimators even when the Gaussian VAR structure assumed by Johansen is true. Following Monte Carlo-based recommendations of Park and Ogaki (1991), the third-step CCR estimates and fourth-step CCR test results are reported.

cointegration is implied. Let *H*(*p, q*) denote the standard Wald statistic to test the hypothesis $\eta_p = \eta_{p+1} = ... = \eta_q = 0$. *H*(*p, q*) converges to a χ^2_{q-p} random variable under the null hypothesis of cointegration, and diverges to infinity under the alternative of no cointegration. With appropriate choice of *p* and *q*, one can test for deterministic and stochastic cointegration. For example, with *p=0* and *q* = 1, the *H*(*0, 1*) statistic tests the deterministic cointegration restriction. With *p=1* and *q=2* or *3*, the *H*(*1, q)* statistics test stochastic cointegration.

Table 2 presents the results for the canonical Balassa-Samuelson model, in which the real exchange rate between two countries is determined solely by the domestic and foreign sectoral *TFP*-differentials as embodied by equation (8). Both the *H*(1, 2) and *H*(1, 3) tests reject the null hypothesis of stochastic cointegration at a 10% level of significance or better for Canada, Denmark, Finland and Norway. For Italy and Sweden, the *H*(1, 2) test statistic is significant at the 10% and 5% levels, respectively, but the *H*(1, 3) test statistic does not reject the null hypothesis of stochastic cointegration. For the remaining 9 countries, both the *H*(1, 2) and *H*(1, 3) tests do not reject the null hypothesis of stochastic cointegration. The *H*(0, 1) test statistic does not reject the deterministic cointegration restriction implied by the Balassa-Samuelson model for any of the countries.

The coefficients of home and foreign sectoral TFP differentials are expected to have positive and negative signs, respectively, and their magnitude is predicted to be between 0 and 1. The coefficient of the home sectoral TFP differential has the right sign for 11 out of 15 countries, and is statistically significant in 8 of these cases. For Australia and China, this coefficient is significantly negative.[10] The coefficient of the U.S. sectoral TFP differential has the predicted negative sign in 9 out of 15 countries and it is statistically significant for 6 of them. This coefficient has the incorrect sign and is statistically significant for 3 of them. The magnitudes of the coefficients which are correctly signed are plausible in most cases after taking into account their standard errors.

Although the results from these cointegration tests are not unambiguous, there is considerable support for the idea that one important factor explaining persistent changes in real exchange rates is different stochastic trends in sectoral TFPs in the home and foreign countries. Figure 1 plots the actual and the implied “equilibrium” values of the bilateral

[10] The negative sign could occur if the home sectoral TFP differential of these countries is cointegrated with the U.S. sectoral TFP differential, thereby rendering the coefficients of the cointegrating regression unidentified.

real exchange rates based on the benchmark Balassa-Samuelson model for the 15 countries in our panel. Despite a few exceptions like Finland, Italy and the Netherlands, for most of the countries the fitted values do seem to capture the long run movements of the real exchange rates.

For example, for the Canadian real exchange rate, the model captures the depreciation of the Canadian dollar from 1975 to 1985, part of the subsequent appreciation from 1985 to 1992, some of the depreciation up to 2002 and the sharp appreciation after 2002.

For France and Germany, the model misses out on the initial appreciation against the U.S. dollar from 1970 to 1980 and the sharp depreciation from 1980 to 1985, but captures quite well the appreciation against the dollar from 1985 to 1995, the subsequent depreciation from 1995 to 2001 as well as the appreciation from 2001 to 2006.

To summarize, the visual evidence of Figure 1 corroborates the statistical results in favor of cointegration for most of the countries in that the long run movements of real exchange rates can be explained for most countries. However, the fitted values are generally unable to capture the medium- and short-run movements of the real exchange rates for several countries.

Table 3 presents the results for the cointegration between real exchange rates and a broader set of fundamentals, including home and foreign sectoral TFP differentials, real interest rate differentials, and the real price of gold, as described in equation (9). The $H(1, 2)$ and $H(1, 3)$ test statistics do not reject the null hypothesis of stochastic cointegration for any of the countries at the 5% level of significance, and are significant only for Japan at the 10% significance level. Moreover, the p-values of $H(1, 2)$ and $H(1, 3)$ test statistics are generally significantly higher for most countries than the corresponding p-values in Table 2, suggesting that the augmented Balassa-Samuelson model better captures the long run relationship between real exchange rates and fundamentals. The $H(0, 1)$ test statistic does not reject the deterministic cointegration restriction implied by the model for any of the countries except Japan, for which it is significant at the 1% level. This is strong evidence in favor of the augmented Balassa-Samuelson model.

The coefficient of the domestic sectoral *TFP* differential has the predicted positive sign for 13 out of 14 countries, and it is statistically significant for all of them. The magnitude of this coefficient is also plausible for these 13 countries. The coefficient of the U.S. sectoral *TFP* differential has the appropriate negative sign for 10 out of 14 countries and is statistically significant for 6 of them. While this coefficient has the wrong sign for 4 countries, it is not statistically significant in 3 of those cases.

The coefficient of the real interest rate differential is positive for 10 countries, and is statistically significant for 9 of them at conventional significance levels. The positive sign is consistent with the evidence documented by Eichenbaum and Evans (1995), Bacchetta and van Wincoop (2010) and Engel (2012).[11] It is significantly negative for Denmark and the UK. The coefficient of the real price of gold has the expected positive sign for all 14 countries, and it is statistically significant for 10 of them at the 1% level of significance.

Figure 2 plots the actual real exchange rates and their fitted values based on the augmented Balassa-Samuelson model. There is a dramatic improvement in the fit of the model for most of the countries, and in addition to the long run movements, the fitted values also explain most of the medium term movements in real exchange rates.

Consider for instance the U.S.−Canadian real exchange rate in the third panel of Figure 2. The model captures the appreciation of the Canadian dollar from 1970 to 1976, the subsequent strong appreciation of the U.S. dollar from 1976 to 1985, the ensuing depreciation of the U.S. dollar from 1985 to 1991, the appreciation of the U.S. dollar from 1991 to 2002, and finally the depreciation of the U.S. dollar from 2002 to the end of the sample in 2006. The model catches most of the turning points of the real exchange rate and matches the volatility of the actual real exchange rate. For most of the countries, the model is able to account for the long swings in real exchange rates reasonably well, although it does miss out on some of the short-run movements.

Indeed, Japan is the only country for which the augmented Balassa-Samuelson model does not perform satisfactorily in explaining the long swings of 1980s and 1990s, and this outcome is rather surprising because Japan is regarded as a classic example of the Balassa-Samuelson effect (Marston, 1987).

Taken together, the evidence presented in Table 3 is quite favorable to the augmented Balassa-Samuelson model. The signs of the coefficients are generally as expected, and the H(p, q) tests do not reject the null hypothesis of stochastic cointegration and the deterministic cointegration restriction for most of the countries. The visual evidence from

[11] Since these papers define the real exchange rate as the relative price of foreign goods to home goods, the real interest rate differential has the opposite (negative sign) in these studies in contrast to the positive sign in ours.

Figure 2, which shows a close link between actual and fitted values of the real exchange rates for most countries, corroborates the statistical evidence.

5 Conclusions

This paper attempts to marry two streams of empirical work on real exchange rates – one emphasizing the importance of real shocks to technology in the tradable and nontradable sectors, and the other focusing on monetary shocks affecting real interest rates and thereby real exchange rates. We also posit a third possible source of real exchange rate movements – the real price of gold serving as a proxy for shocks to the financial system. In contrast with the bulk of the previous literature, which generally finds that macroeconomic economic fundamentals have weak explanatory power for persistent departures of exchange rates from PPP, this approach yields both economically and statistically greater explanatory power.

These results are also consistent with some recent work that emphasizes the dichotomy between tradable and nontradable goods in explaining long run movements in real exchange rates. (e.g., Burstein, Eichenbaum and Rebelo, (2005a, 2005b), Betts and Kehoe (2006), Crucini and Shintani (2008), Engel (2000) and Park and Ogaki (2007)).

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Tables and figures

Table 1a Unit root tests[1]

	Log real exchange rate ($ln(RER)$)			Log productivity differential (d)			Long-term interest rate differential ($r^{LT\ dif}$)		
	J(1,5)[3]	Phillips-Perron[4] (no trend)	Phillips-Perron[4] (trend)	J(1,5)[3]	Phillips-Perron[4] (no trend)	Phillips-Perron[4] (trend)	J(1,5)[3]	Phillips-Perron[4] (no trend)	Phillips-Perron[4] (trend)
Australia	1.337	−2.241	−2.640	1.064	−2.290	−2.722	1.169	−2.483	−2.332
Belgium	0.775	−2.517	2.462	13.292	−2.140	−1.048	0.206**	−2.608	−3.274*
Canada	2.316	−1.833	−2.015	3.084	−1.477	−2.544	0.604	−2.053	−2.036
Denmark	0.945	−2.526	−2.495	1.596	−1.921	−2.324	0.826	−0.880	−3.619**
Finland	1.055	−2.679*	−2.490	2.574	−1.833	−2.067	7.447	−0.875	−2.731
France	0.622	−2.469	−2.378	5.404	−1.365	−1.976	0.133**	−2.281	−4.423***
Germany	0.986	−2.480	−2.311	4.833	−0.304	−2.274	4.419	−1.836	−1.749
Italy	0.728	−1.961	−2.477	2.676	−1.684	−2.019	1.215	−1.666	−2.067
Japan	2.232	−2.463	−1.532	1.164	−1.636	−1.250	1.307	−1.770	−2.145
S. Korea	1.080	−1.954	1.968	0.095***	1.842	−2.872	1.379	−1.253	−2.211
Netherlands	0.958	−2.671*	−2.523	5.818	−0.836	−1.377	1.594	−2.100	−2.060
Norway	1.628	−2.139	−2.246	0.522	−1.974	−2.029	0.612	−2.417	−2.294
Sweden	0.734	−2.446	−2.477	0.086***	−1.667	−2.411	1.289	−2.027	−1.903
UK	0.127**	−1.751	−2.630	4.420	−1.123	−0.851	1.145	−1.812	−2.703
US	-	-	-	6.329	−1.019	−2.107	-	-	-

Notes:

1. The null hypothesis is the existence of unit root.
2. "***" indicates significance at 1%, "**" indicates significance at 5% and "*" indicates significance at 10%.
3. $J(1, 5)$ denotes Park's (1990) test for the null hypothesis of a unit root against the alternative of trend stationarity.
4. The **Phillips-Perron** (PP) test refers to the test developed by Phillips and Perron (1988).

Table 1b Unit root tests1 (J(1,5) and Phillips-Perron tests)

	Log real gold price ($ln(gold)$)		
	J(1,5)	Phillips-Perron[3] (no trend)	Phillips-Perron[3] (trend)
World	4.3218	−2.8657*	−2.8923

Notes:

1. The null hypothesis is the existence of unit root.
2. "***" indicates significance at 1%, "**" indicates significance at 5% and "*" indicates significance at 10%.
3. $J(1, 5)$ denotes Park's (1990) test for the null hypothesis of a unit root against the alternative of trend stationarity.

Table 2 CCR regressions with productivity differentials

$$\ln(E^r) = \theta + \beta^H d^H - \beta^F d^F + \varepsilon$$

		Coefficient estimates[1]			Cointegration tests[2,3]		
Country	Sample	Constant	d_H	d_{US}	H(0,1)	H(1,2)	H(1,3)
Australia	1970–2006	0.689*** (0.123)	−0.753** (0.306)	−1.009*** (0.323)	0.201 [0.654]	1.104 [0.293]	2.832 [0.243]
Belgium	1970–2006	−2.976*** (0.181)	1.152*** (0.315)	−3.177*** (0.818)	2.323 [0.127]	0.761 [0.383]	1.302 [0.522]
Canada	1970–2006	0.243*** (0.073)	0.163** (0.074)	−0.762*** (0.196)	0.928 [0.335]	2.920 [0.088]*	7.662 [0.022]**
Denmark	1970–2006	−1.604*** (0.055)	0.812*** (0.081)	−1.022*** (0.176)	1.306 [0.253]	4.529 [0.033]**	14.303 [0.000]***
Finland	1970–2006	−1.114*** (0.130)	0.202* (0.122)	0.246 (0.339)	0.389 [0.533]	4.522 [0.034]**	9.489 [0.009]***
France	1970–2006	−1.089*** (0.141)	0.547*** (0.157)	−1.399*** (0.430)	0.328 [0.567]	1.720 [0.190]	1.730 [0.421]
Germany	1970–2006	−0.854*** (0.103)	0.850*** (0.176)	−2.109*** (0.471)	0.182 [0.670]	0.027 [0.871]	2.803 [0.246]
Italy	1970–2006	−6.274*** (0.107)	−0.071 (0.240)	0.523** (0.246)	0.015 [0.903]	3.279 [0.070]*	3.309 [0.191]
Japan	1970–2006	−6.580*** (0.693)	0.415 (0.290)	0.344 (0.601)	0.379 [0.538]	0.403 [0.525]	0.502 [0.778]
S. Korea	1970–2006	−5.041*** (0.222)	0.337*** (0.117)	−0.301 (0.574)	0.086 [0.769]	0.774 [0.379]	0.843 [0.656]
Netherlands	1970–2006	−0.974*** (0.133)	0.040 (0.088)	0.322 (0.356)	0.556 [0.456]	3.191 [0.074]*	3.253 [0.197]
Norway	1970–2006	−1.538*** (0.112)	−0.040 (0.055)	−0.301 (0.287)	2.695 [0.101]	5.081 [0.024]*	10.529 [0.005]***
Sweden	1970–2006	−2.063*** (0.195)	0.905*** (0.215)	0.729** (0.363)	1.708 [0.191]	4.440 [0.035]**	4.499 [0.105]
UK	1970–2006	0.735*** (0.194)	0.154 (0.219)	1.399*** (0.504)	0.159 [0.691]	0.022 [0.883]	0.242 [0.886]

Notes:
1. For the coefficient estimates, standard errors are in parentheses.
2. For the cointegration tests, the chi-square test statistics are reported. P-values are in the square bracket.
3. H(0,1) is for testing the deterministic cointegrating restriction. H(1,2) and H(1,3) are for testing the stochastic cointegrating restriction. In these H(p,q) tests, q refers to the order of polynomial in the added time trends.
4. "***" indicates significance at 1%, "**" indicates significance at 5% and "*" indicates significance at 10%.

Table 3 CCR estimation (with productivity differentials, long-term interest rate differential and ln(real gold price))

$$\ln(E^r) = \theta + \beta^H d^H - \beta^F d^F + \beta_r r_{diff} + \beta_g \ln(rp_{gold}) + \varepsilon',$$

		Coefficient estimates[1]					Cointegration tests[2,3]		
Country	Sample	Constant	d_H	d_{US}	$r^{LT\ dif}$	$ln(real\ gold)$	H(0,1)	H(1,2)	H(1,3)
Australia	1970−2006	−0.914*** (0.184)	0.283 (0.126)	−0.643*** (0.128)	0.019* (0.011)	0.301*** (0.036)	1.699 [0.192]	0.080 [0.778]	1.766 [0.414]
Belgium	1970−2006	−3.788*** (0.348)	0.628*** (0.137)	−2.045*** (0.409)	0.195*** (0.051)	0.121 (s.e.= 0.074)	0.567 [0.451]	0.004 [0.950]	0.180 [0.914]
Canada	1970−2006	−0.806*** (0.159)	0.259*** (0.090)	−0.987*** (0.140)	0.016 (0.033)	0.235*** (0.043)	0.020 [0.887]	0.024 [0.877]	1.128 [0.569]
Denmark	1970−2006	−2.623*** (0.196)	0.685*** (0.067)	−1.243*** (0.188)	−0.030** (0.012)	0.257*** (0.052)	0.676 [0.411]	0.039 [0.845]	0.040 [0.980]
Finland	1988−2006	−3.057*** (0.275)	0.110*** (0.034)	−0.018 (0.157)	0.022*** (0.004)	0.437*** (0.045)	1.6679 [0.1965]	2.5260 [0.1120]	2.7151 [0.2573]
France	1970−2006	−1.764*** (0.295)	0.524*** (0.109)	−0.602** (0.284)	0.135*** (0.049)	0.068 (0.068)	0.793 [0.373]	0.984 [0.321]	0.986 [0.611]
Germany	1970−2006	−3.003*** (0.312)	0.362*** (0.111)	0.021 (0.312)	0.117*** (0.020)	0.394*** (0.066)	1.450 [0.229]	0.101 [0.750]	0.622 [0.733]
Italy	1970−2006	−6.860*** (0.343)	0.454*** (0.162)	0.559** (0.234)	0.092*** (0.027)	0.060 (0.073)	0.4089 [0.5225]	2.3903 [0.1221]	2.8142 [0.2449]
Japan	1970−2006	−7.222*** (0.849)	0.498* (0.297)	−0.622 (0.934)	−0.032 (0.054)	0.140 (0.133)	1.931 [0.165]	3.781 [0.052]*	5.432 [0.066]*
S. Korea	1974−2006	−7.372*** (0.326)	0.380*** (0.037)	0.072 (0.178)	0.066*** (0.008)	0.390*** (0.067)	0.786 [0.375]	1.163 [0.281]	2.840 [0.242]
Nether.	1970−2006	−3.378*** (0.225)	0.307 *** (0.055)	−0.221 (0.189)	0.069*** (0.016)	0.5569*** (0.053)	0.203 [0.653]	0.031 [0.859]	0.222 [0.895]
Norway	1970−2006	−2.361*** (0.248)	−0.128*** (0.041)	−0.851*** (0.274)	−0.006 (0.019)	0.219*** (0.055)	10.890 [0.001]***	1.235 [0.266]	1.260 [0.533]
Sweden	1970−2006	−2.154*** (0.230)	0.313*** (0.090)	−0.228 (0.177)	0.066*** (0.015)	0.127*** (0.047)	2.023 [0.155]	0.016 [0.900]	0.574 [0.751]
UK	1970−2006	−0.362 (0.241)	0.613*** (0.081)	0.079 (0.198)	−0.062** (0.016)	0.322*** (0.054)	0.487 [0.485]	0.117 [0.732]	0.343 [0.843]

Notes:

1. For the coefficient estimates, standard errors are in parentheses.
2. For the cointegration tests, the chi-square test statistics are reported. P-values are in the square bracket.
3. H(0,1) is for testing the deterministic cointegrating restriction. H(1,2) and H(1,3) are for testing the stochastic cointegrating restriction. In these H(p,q) tests, q refers to the order of polynomial in the added time trends.
4. "***" indicates significance at 1%, "**" indicates significance at 5% and "*" indicates significance at 10%.

Figure 1 Actual and equilibrium real exchange rates based on productivity differentials

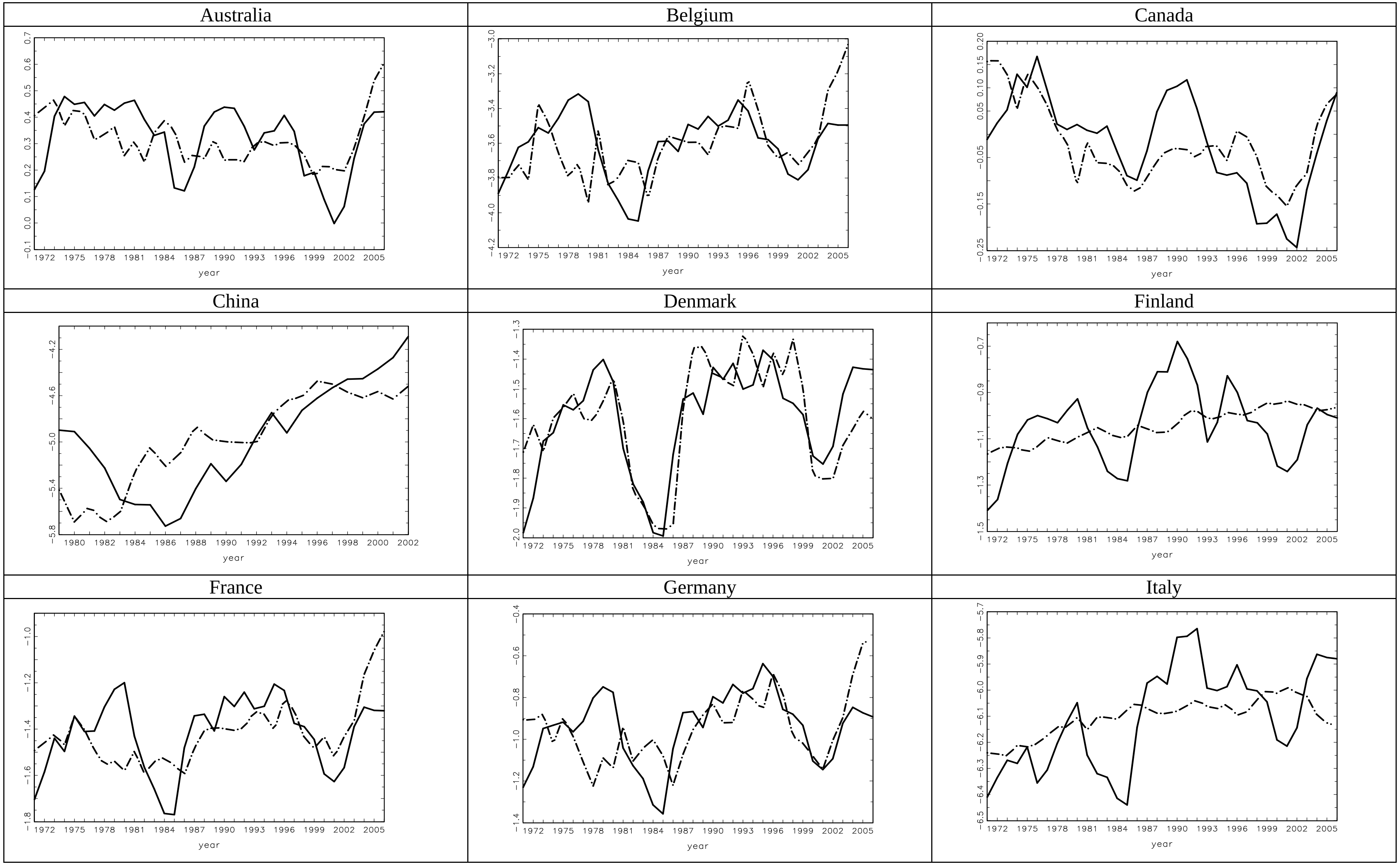


Figure 1 Actual and equilibrium real exchange rates based on productivity differentials (continued)

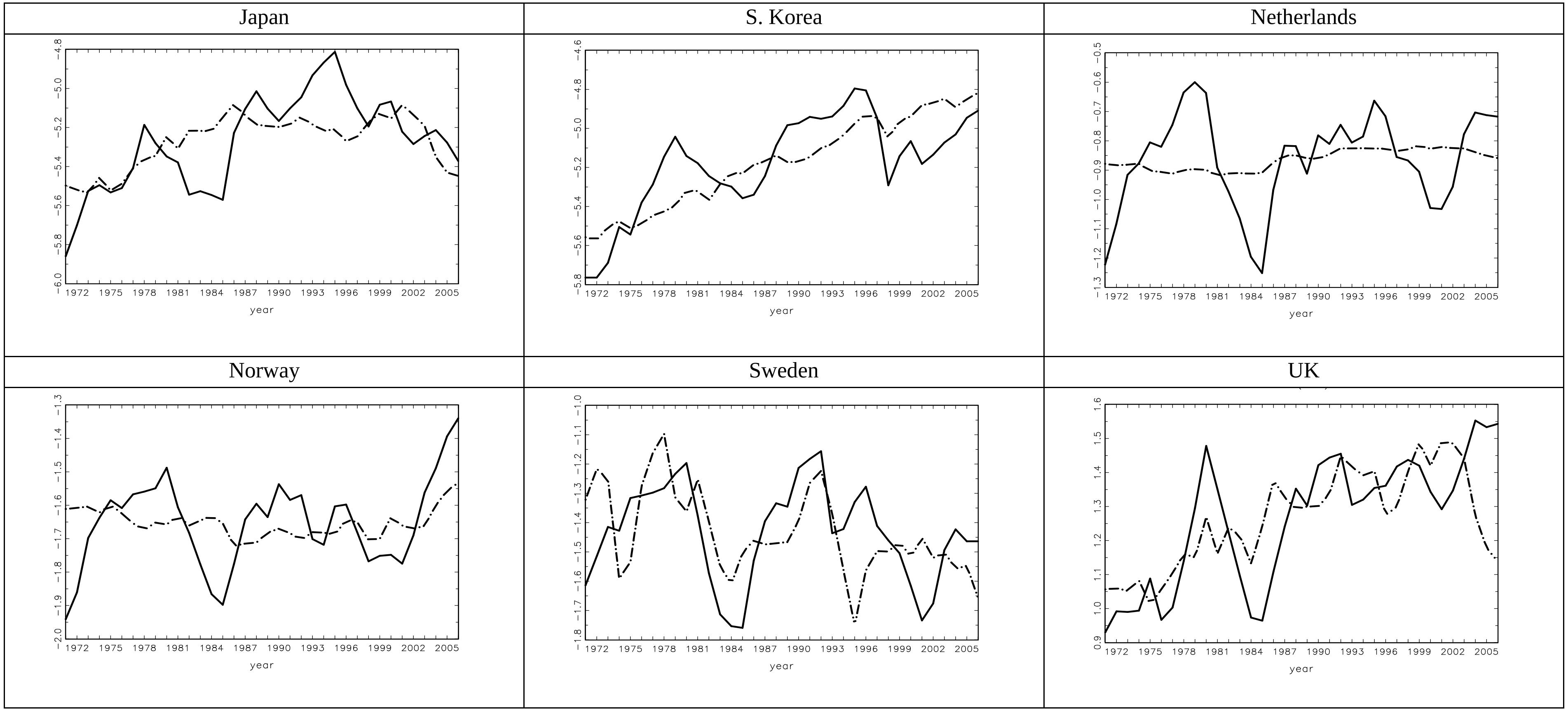


Figure 2 Actual and equilibrium real exchange rates based productivity differentials, long-term interest rate differential and real gold price

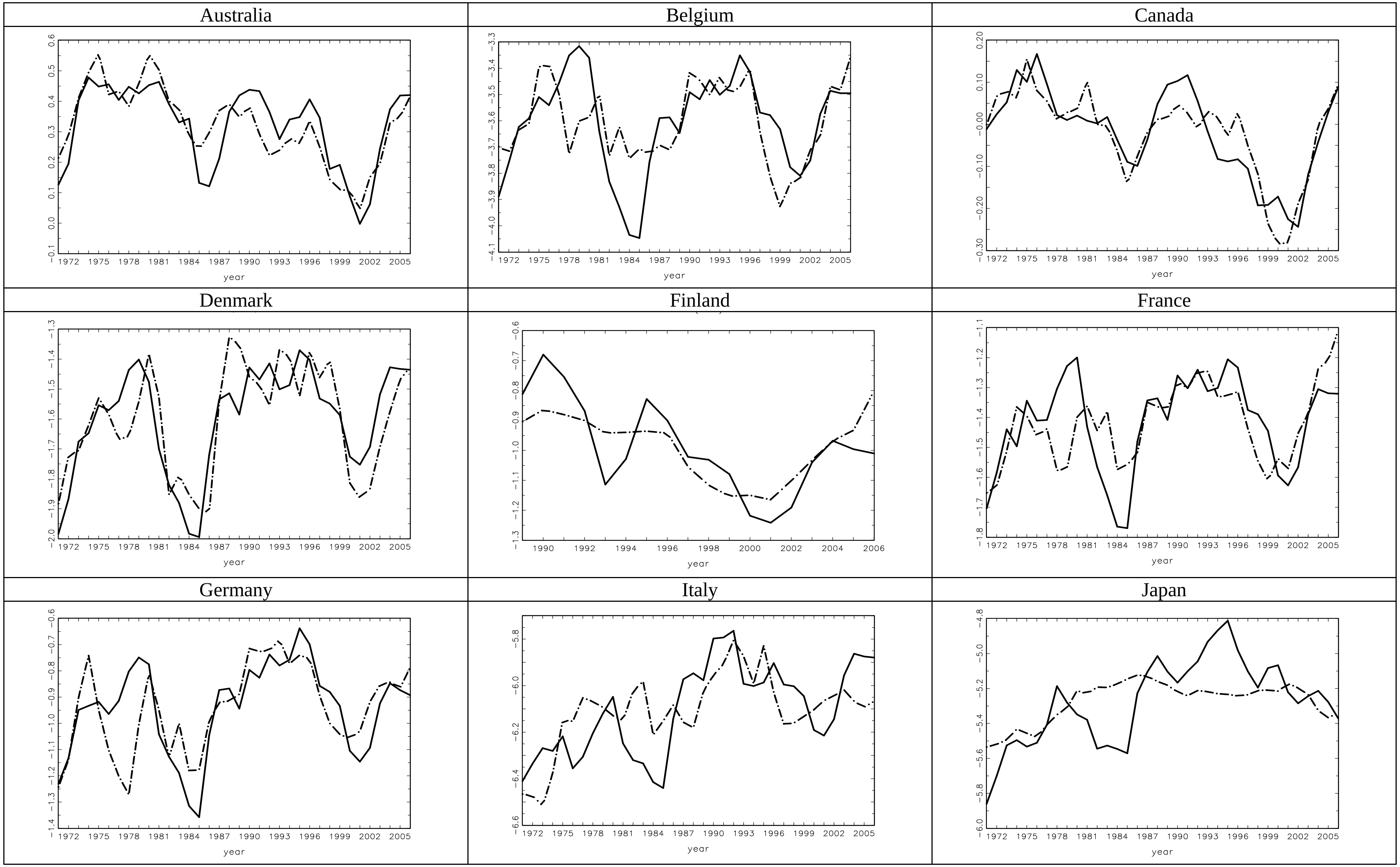
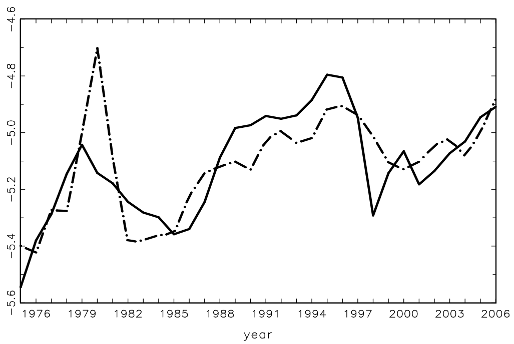
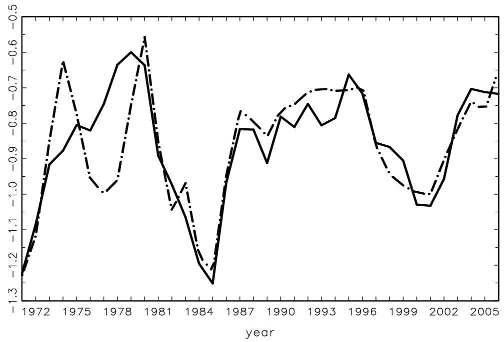


Figure 2 Actual and equilibrium real exchange rates based productivity differentials, long-term interest rate differential and real gold price (continued)

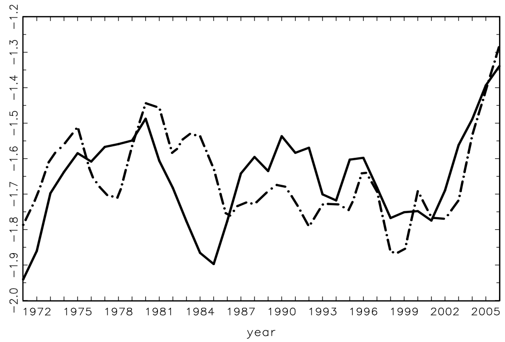
S.Korea



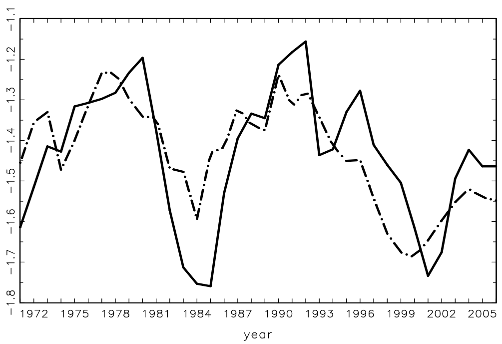
Netherlands



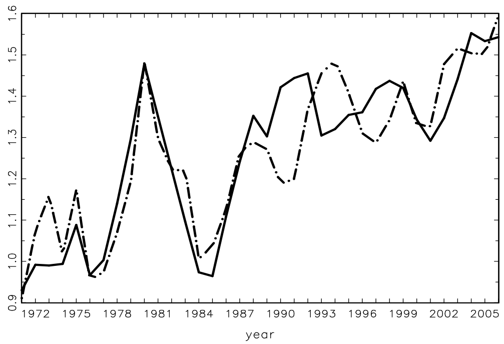
Norway



Sweden



UK



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