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## Conventional and Unconventional Approaches to Exchange Rate Modeling and Assessment

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# Conventional and Unconventional Approaches to Exchange Rate Modeling and Assessment

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## **Abstract**

We examine the relative predictive power of the sticky price monetary model, uncovered interest parity, and a transformation of the net exports variable. In addition to bringing a new approach (utilizing our measure of external imbalance suggested by Gourinchas and Rey) and data spanning a more recent period to bear, we implement the Clark and West (forthcoming) procedure for testing the significance of out-of-sample forecasts. The interest rate parity relation holds better at long horizons and the net exports variable does well in predicting exchange rates at short horizons in-sample. In out-of-sample forecasts, we find evidence that uncovered interest parity outperforms a random walk at long horizons and that the measure of external imbalance does well at short horizons, although we cannot duplicate the findings of Gourinchas and Rey.

**Key words:** exchange rates, monetary model, net foreign assets, interest rate parity, forecasting performance

**JEL classification:** F31, F47

## 1. Introduction

Over the last two years, movements in dollar exchange rates have proven as inexplicable as ever. We take up the issue of whether there are enduring and significant instances where economic models can explain and predict future movements in exchange rates.

First, we examine the behavior of several key dollar exchange rates during the first euro/dollar cycle. There is ample interest in the behavior of the euro, aside from the synthetic euro, which provided the basis of previous studies, as in Chinn and Alquist (2000) and Schnatz et al. (2004). Second, we examine the relative performance of a model incorporating a role for net foreign assets, as suggested in Gourinchas and Rey's (2005) financial adjustment channel as well as Lane and Milesi-Ferretti (2003). Third, we employ a new test that is appropriate for testing nested models. Standard tests for the statistical significance of out-of-sample forecasts are wrongly sized.<sup>1</sup>

Gourinchas and Rey recently suggested that multilateral dollar exchange rates are well predicted by a procedure that takes into account financial adjustment. The channel is the natural implication of an intertemporal budget constraint that allows for valuation changes in foreign assets and liabilities. More broadly, there is a large literature that links foreign assets and liabilities to exchange rates (Lane and Milesi-Ferretti, 2005). The examination provides an opportunity to determine if the finding is replicable using alternative data sets, different sample periods, and different currencies.

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<sup>1</sup> The evaluation procedure also differs slightly from that in Cheung et al. (2005a,b). We evaluate the mean squared prediction error of the predicted change in the exchange rate rather than the level. In the context of the error correction forecasts we make, and conditioning on the current spot exchange rate, the resulting comparison is the same; to stay close to the spirit of the Clark and West (forthcoming) paper, however, we retain the comparison in changes.

In a recent paper, Engel and West (2005) have argued that one should not expect much exchange rate predictability, given that the fundamentals for exchange rates are highly persistent processes and the discount factor follows very closely a random walk. We do not dispute the view, especially given the negative results in Cheung et al.'s (2005a,b) comprehensive study of several competing models.<sup>2</sup> The study, however, did not deploy the most appropriate test. In the current paper study, we remedy the deficiency by using Clark and West (forthcoming) method.<sup>3</sup>

We summarize the exchange rate models considered in the exercise in Section 2. Section 3 discusses the data and in-sample fit of the models. Section 4 outlines the forecasting exercise, estimation methods, and the criteria used to compare forecasting performance. The forecasting results are reported in Section 5. Section 6 concludes.

## 2. The Models and Some Evidence

We use the random walk model as our benchmark model, in line with previous work. As the workhorse model, we appeal to the Frankel (1979) formulation of the Dornbusch (1976) model, as it provides the fundamental intuition for how flexible exchange rates behave. The sticky price monetary model is:

$$(1) \quad s_t = \beta_0 + \beta_1 \hat{m}_t + \beta_2 \hat{y}_t + \beta_3 \hat{i}_t + \beta_4 \hat{\pi}_t + u_t,$$

where  $m$  is log money,  $y$  is log real GDP,  $i$  and  $\pi$  are the interest and inflation rate, and  $u_t$

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<sup>2</sup> See Faust et al. (2003), while MacDonald and Marsh (1999), Groen (2000) and Mark and Sul (2001) provide more positive results.

<sup>3</sup> Other papers using the Clark-West approach include Gourinchas and Rey (2005) and Molodtsova and Papell (2006).

is an error term. The circumflexes denote inter-country differentials. The characteristics of the model are well known. The money stock and the inflation rate coefficients should be positive; and the income and interest rate coefficients negative, as long as prices are sticky. If prices are perfectly flexible, either interest rates or inflation rates should enter in positively.

The specification is more general than it appears. The variables included in (1) encompass the flexible price version of the monetary model, as well as the micro-based general equilibrium models of Stockman (1980) and Lucas (1982). In addition, the sticky price model is an extension of equation (1), where the price variables are replaced by macro variables that capture money demand and overshooting effects. We do not impose coefficient restrictions in equation (1) because theory provides little guidance regarding the values of the parameters.

The next specification assessed is not a model *per se*; rather it is an arbitrage relationship – uncovered interest rate parity<sup>4</sup>:

$$(2) \quad s_{t+k} - s_t = \hat{i}_t^k,$$

where  $\hat{i}_t^k$  is the interest rate of maturity  $k$ .<sup>5</sup> This relationship need not be estimated to generate predictions.

The interest rate parity relationship is included in the forecast comparison

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<sup>4</sup> To be technically correct, (2) represents uncovered interest parity combined with unbiased expectations. This is sometimes termed the unbiasedness hypothesis.

<sup>5</sup> For notational consistency, we use the log approximation in discussing interest rate parity, but use the exact expression in the regressions and the forecasting exercises. The results do differ somewhat between the two methods, particularly when the sample includes the 1970s when interest rates were relatively high.

exercise because it has recently gathered empirical support at long horizons (Chinn and Meredith, 2004), in contrast to the disappointing results at the shorter horizons. Cheung et al. (2005a,b) confirm that long-run interest rates predict exchange rate levels better than alternative models.<sup>6</sup>

The third model is based upon Gourinchas-Rey (2005). Gourinchas and Rey (2005) log-linearize a transformation of the net exports to net foreign assets variable around its steady state value. Appealing to the long run restrictions imposed by assumptions of stationarity of asset to wealth, liability to wealth, and asset to liability ratios, they find that either trade flows, portfolio returns, or both, adjust. By way of contrast, the traditional intertemporal budget approach to the current account takes trade flows as the principal object of interest, probably because most of the models in this vein contain only one good.

The financial channel implies that the net portfolio return,  $ret$ , which combines market and exchange rate induced valuation effects, exhibits the following relationship:

$$(3) \quad ret_t = \gamma + \lambda nxa_{t-1} + Z_{t-1}\Xi + \varepsilon_t$$

where  $Z$  is a set of control variables, and

$$(4) \quad nxa_t = \frac{\mu_m}{\mu_x} xm_t - \frac{\mu_l}{\mu_x} al_t + \frac{1}{\mu_x} xa_t$$

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<sup>6</sup> Despite the finding, there is little evidence that long-term interest rate differentials – or, equivalently, long-dated forward rates – have been used for forecasting at the horizons we are investigating. One exception from the non-academic literature is Rosenberg (2001).

where the  $\mu$ 's are normalized weights;  $xm$  is the log export to import ratio;  $al$  is the log asset to liability ratio; and  $xa$  is the log export to asset ratio. When one normalizes the exports variable to have a coefficient of unity as in (4), Gourinchas and Rey describe the  $nxa$  variable as: "approximately the percentage increase in exports necessary to restore external balance (i.e., compensate for the deviation from trend of the net exports to net foreign asset ratio)." (2005, page 12).

They conclude that the financial adjustment channel is quantitatively important at the medium frequency. That is, asset returns do fair bit of the adjustment at horizons of up to two years; thereafter, the conventional trade balance channel takes effect. Exchange rates, as part of the financial adjustment process as well as the trade balance process, occupy a dual role. They find that at the one quarter ahead horizon, 11% of the variance of exchange rate is predicted, while at one and three years ahead, 44% and 61% of the variance is explained. Specifically, they statistically outperform a random walk at all horizons between one to twelve quarters, using the Clark and West (forthcoming) test method.<sup>7</sup>

### **3. Data and In-sample Model Fit**

To provide some insight into how plausible the models are, we conduct some regression analysis to show whether the specifications make sense, based on in-sample diagnostics.<sup>8</sup>

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<sup>7</sup> They use an initial estimation window of 1952q1-1978q1, and roll the regressions for the cointegrating vector.

<sup>8</sup> For a discussion of why one might want to rely solely on in-sample diagnostics, see Inoue and Kilian (2004).

### **3.1 Data**

We rely upon quarterly data for the United States, Canada, U.K., and the Euro Area over the 1970q1 to 2004q4 period.<sup>9</sup> The exchange rate, money, price and income (real GDP) variables are drawn primarily from the IMF's *International Financial Statistics*. M1 is used for the money variable, with the exception of the UK, where M4 is used. For the money stocks, exchange rates, and interest rates, end-of-quarter rates are used. The interest rate data used for the interest rate parity estimates are from the national central banks as well as Chinn and Meredith (2004). The Euro Area data are drawn from the Area Wide Model, described in Fagan et al. (2001). The end of year U.S. foreign asset and liability data are from the Bureau of Economic Analysis (BEA) and interpolated using quarterly financial account data from *IFS*.

The measure of external imbalance requires some discussion. The central point made by Gourinchas and Rey (2005) and Lane and Milesi-Ferretti (2005) is that cumulated trade balances can be very inaccurate measures of net foreign asset positions in an era of financial integration, i.e., where gross asset and liability positions are large relative to GDP, and are subject to substantial variation due to exchange rate induced valuation changes. BEA provides end-of-year data on gross asset and gross liability positions, denominated in dollars. We generate quarterly data by distributing measured financial account balance data over the year.<sup>10</sup> The Data Appendix contains more detail.

### **3.2 Estimation and Results**

Using DOLS, we check the monetary model for a long-run relationship between the

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<sup>9</sup> Data series for the euro area begin around 1980.

<sup>10</sup> In the previous version of this paper, we used the Lane and Milesi-Ferretti data.

exchange rate and the explanatory variables (Stock and Watson, 1993):

$$(5) \quad s_t = X_t \Gamma + \sum_{i=-2}^{+2} \Delta X_t B + u_t,$$

where  $X$  is the vector of explanatory variables associated with the monetary approach.

For the other models, the specifications take on an “error-correction”-like form. The interest rate parity relationship in (2) implies the annualized change in the exchange rate equals the interest rate differential of the appropriate maturity. The Gourinchas-Rey specification of equation (3) implies that exchange rate changes are related the lagged level of the  $nx_a$ . In these instances, we rely upon standard OLS estimates, relying upon heteroskedasticity and serial correlation robust standard errors to conduct inference.<sup>11</sup>

Table 1 displays the estimated cointegrating vectors for the monetary model normalized on the exchange rate. Theory predicts that money enters with a positive coefficient, income with a negative coefficient, and inflation with a positive coefficient. Interest rates have a (negative) independent effect above and beyond that of inflation if prices are sticky, so that higher nominal interest rates, holding inflation constant, appreciate the home currency.

As usual, the estimated money coefficients are wrongly signed and insignificant. Money has the right sign for Canada, but that is the sole case; in any event, the value is substantially smaller than suggested by theory. Income and interest rates typically have

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<sup>11</sup> For the  $nx_a$  regressions, we follow the suggestion of Newey and West and set the truncation lag equal to  $4(T/100)^{2/9}$ . In the interest rate parity regressions, where under the unbiasedness null the errors follow an MA(k-1) process, we use a truncation lag of  $2(k-1)$  indicated by Cochrane (1991).

the right sign, but inflation appears to exert a more consistently significant effect. Hence, we have some slight evidence in favor of a long run cointegrating relationship between monetary fundamentals and nominal exchange rates in the three cases.

The interest rate parity results in Table 2 replicate those found in Chinn and Meredith (2004): at short horizons, the coefficient on interest rates point in a direction inconsistent with the joint null hypothesis of uncovered interest parity and unbiased expectations (note that the significance levels in this panel are for the null of the slope coefficient equaling unity).

For the euro/dollar rate, we report results for a sample that pertains to the post-EMU sample, and a longer one (back to 1990) relying upon the synthetic euro, and using the German 3 month rate as a proxy for the euro interest rate. In both cases, the forward discount points in the wrong direction, on average, with the bias more pronounced in the shorter post-EMU period.

At the long horizon of 5 years, interest rates point in the right direction, and the null of a unitary coefficient cannot be rejected at conventional levels. This pattern of results matches that reported in Cheung et al. (2005b), as well as Chinn and Meredith (2004). While the results are promising, the adjusted  $R^2$ 's are still quite low.

Finally, we turn to the regression with our measure of external imbalance. The series we have generated using BEA data is plotted in Figure 2. We cannot exactly replicate the patterns in the Gourinchas-Rey series, which is not surprising as we do not undertake as detailed a reconstruction of the underlying assets and liabilities.

Theory predicts the exchange rate depreciates in response to a negative value of the explanatory variable. Given the way in which the exchange rate is defined, this

implies a negative sign. Table 3 reports some positive results, in that the coefficients are statistically significant in the right direction 8 out of 9 times.

In addition, the effect appears more pronounced at the short and short to intermediate horizon (3 month, 1 year) than at 5 years. In addition, the adjusted  $R^2$ 's for these regressions is fairly high by comparison to the long horizon interest rate regressions. These results augur well for positive results on the forecasting end.

#### **4. Forecasting Procedure and Comparison**

##### ***4.1 The Forecasting Exercise***

To ensure that the conclusions are not sensitive to the choice of a specific out-of-forecasting period, we use two out-of-sample simulation periods to assess model performance: 1987q2-2004q4 and 1999q1-2004q4. The former period conforms to the post-Louvre Accord period, while the latter spans the post-EMU period. The longer out-of-sample period (1987-2004) spans a period of *relative* dollar stability with one upswing and downswing in the dollar's value.

We adopt the convention in empirical exchange rate modeling of implementing “rolling regressions” established by Meese and Rogoff (1983). We estimate the model for a given sample, use the parameter estimates to generate out-of-sample forecasts, roll the sample forward one observation, and repeat the procedure. We continue until we exhaust all of the out-of-sample observations. While rolling regressions do not incorporate possible efficiency gains as the sample moves forward through time, the procedure alleviates parameter instability, which is common in exchange rate modeling.

We use an error correction specification for the sticky price monetary model,

while the competing models intrinsically possess an error correction nature. Returning to the monetary model, both the exchange rate and its economic determinants are I(1). The error correction specification allows for the long-run interaction effect of these variables, captured by the error correction term, in generating forecasts. If the variables are cointegrated, then the former specification is more efficient than the latter one and is expected to forecast better in long horizons. If the variables are not cointegrated, the error correction specification leads to spurious results. Since we detect evidence in favor of cointegration, we rely upon error correction specifications for the monetary model.

The general expression for the relationship between the exchange rate and fundamentals:

$$(6) \quad s_t = X_t \Gamma + \varepsilon_t,$$

where  $X_t$  is a vector of fundamental variables indicated in (1). The error correction estimation is a two-step procedure. In the first step, we identify the long-run cointegrating relation implied by (6) using DOLS. Next we incorporate the estimated cointegrating vector ( $\tilde{\Gamma}$ ) into the error correction term, and estimate the resulting equation

$$(7) \quad s_t - s_{t-k} = \delta_0 + \delta_1 (s_{t-k} - X_{t-k} \tilde{\Gamma}) + u_t$$

using OLS. Equation (7) is an error correction model stripped of short run dynamics.

Mark (1995) and Chinn and Meese (1995) use a similar approach, except that they impose the cointegrating vector *a priori*. The specification is motivated by the difficulty

in estimating the short run dynamics in exchange rate equations.<sup>12</sup> In contrast to other studies, our estimates of the *long-run* cointegrating relationship vary as the data window moves.<sup>13</sup>

#### 4.2 Forecast Comparison

To evaluate the forecasting accuracy of the different structural models, we use the adjusted-mean squared prediction error (MSPE) statistic proposed by Clark and West (forthcoming). Under the null hypothesis, the MSPE of a zero mean process is the same as the MSPE of the linear alternative. Despite the equality, one expects the alternative model's sample MSPE to be larger than the null's. To adjust the downward bias, Clark and West propose a procedure that performs well in simulations (Clark and West, 2005a,b).

The test statistic is the difference between the MSPE of the random walk model and the MSPE from the linear alternative, which is adjusted downward to account for the spurious in-sample fit. Under the first model, the process is a zero mean martingale difference process; under the second model, the process is linear,

$$(8) \quad \begin{aligned} \text{Model 1: } & y_{t+1} = e_{t+1} \\ \text{Model 2: } & y_{t+1} = X_{t+1}B + e_{t+1}, E(e_{t+1} | I_t) = 0 \end{aligned}$$

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<sup>12</sup> We exclude short-run dynamics in equation (7) for two reasons. First is that the use of equation (7) yields true *ex ante* forecasts and makes our exercise directly comparable with, for example, Mark (1995), Chinn and Meese (1995) and Groen (2000). Second, the inclusion of short-run dynamics creates additional demands on the generation of the right-hand-side variables and the stability of the short-run dynamics that complicate the forecast comparison exercise beyond a manageable level.

<sup>13</sup> We do not impose restrictions on the  $\beta$ -parameters in (1) because we do not have strong priors on the exact values of the coefficients.

Our inferences are based on a formal test for the null hypothesis of no difference in the accuracy (i.e., in the MSPE) of the two competing forecasts, the linear (structural) model and the driftless random walk. Thus, the hypothesis test is

$$H_0 : \sigma_1^2 - \sigma_2^2 = 0$$

$$H_A : \sigma_1^2 - \sigma_2^2 > 0$$

A value larger (smaller) than zero indicates the linear model (random walk) outperforms the random walk (linear model).

The difference between the two MSPEs is asymptotically normally distributed.<sup>14</sup> For forecast horizons beyond one period, one needs to account for the autocorrelation induced by the rolling regression. We use Clark and West's proposed estimator for the asymptotic variance of the adjusted mean between the two MSPEs, which is robust to the serial correlation (Clark and West, forthcoming).

## 5. Comparing the Forecasts

Table 4 presents the results of comparing the forecasts. The top entry in each cell is the difference in the MSPEs (positive entries denote out-performance of a random walk). The Clark-West statistic is displayed below (Diebold-Mariano/West statistic for the interest rate parity results); this can be read as a  $z$ -statistic.

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<sup>14</sup> Since the interest rate parity coefficient of unity is imposed, rather than estimated, we used the Diebold-Mariano (1995)/West (1996) test statistic, assuming the  $h$ -step ahead forecast error follow an MA( $h-1$ ) process.

At most horizons, the interest rate parity model does as well as the random walk.<sup>15</sup> During the 1999q1-2004q4 sample, forecasts based on the interest parity relationship significantly outperform the random walk at the 20-quarter horizon for the Canadian dollar and the euro, although the result does not hold over the longer sample, 1987q2-2004q4. At shorter horizons, the interest rate parity forecast does less well than the random walk forecast. The result is consistent with Cheung et al. (2005a,b), even though we omit the yen, a currency for which interest rate parity substantially outperformed a random walk.

The sticky price monetary model outperforms the random walk in 4-quarter ahead forecasts during 1987q2-2004q4 for the pound, but significantly worse for the Canadian dollar. In the most recent 1999-2004 period, it does worse, and significantly so, at one year horizons. The finding is consistent with Cheung et al. (2005a,b), who point out that the papers in which the sticky-price monetary model outperform the random walk in out-of-sample forecasts usually involve a cointegrating vector estimated over the entire sample, as in MacDonald and Taylor (1994). Our test is more stringent. We estimate the cointegrating vector over a rolling window rather than the entire sample, making the forecasts true ex-ante predictions.

We find favorable support for the use of the measure of external imbalance, particularly when we estimate the cointegrating vector over the longer sample. In the second subsample, it outperforms the random walk at short horizons for the Canadian dollar and the pound at 5 and 10 % level significance level. The results are less favorable

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<sup>15</sup> As opposed to using the interest differentials as fundamentals as in Clark and West (forthcoming) and Molodtsova and Papell (2006), we use the exact formula implicit in the unbiasedness hypothesis.

at long horizons and not particularly positive for the euro.

Both of the results are consistent with those obtained by in Gourinchas and Rey (2005), although our findings are weaker than theirs. Indeed, they find that the lagged measure of external imbalance outperforms the random walk at both short- and long-horizons.

Although our method mimics theirs, there are plausible reasons for the discrepancies. First, the Gourinchas and Rey data are not publicly available, so we constructed our own measure of *nxa* that only approximates theirs. A second, related reason is that Gourinchas and Rey have a time series extending back to the 1950s, which enables them to estimate the cointegrating vector with a sample of 105 quarterly observations. Indeed, Gourinchas and Rey state that they need the long estimation window to obtain stable estimates of the cointegrating vector (Gourinchas and Rey, 2005). By way of contrast, our *nxa* measure begins in 1977q1, giving us an estimation window of 41 observations for the first subsample (1977q1-1987q1) and 88 for the second subsample (1977q1-1998q4). Thus, the differences are likely due to the alternative samples. In fact, the difference in the MSPEs is positive, if not statistically significant, in the sample with 88 observations, suggesting that forecasts based on a longer time series may perform better for bilateral exchange rates.

The finding is surprising, because the principal prediction of the Gourinchas-Rey approach pertains to multilateral exchange rate returns rather than to bilateral exchange rates. According to the model, one should not expect the measure of external imbalance to have predictive power for currencies with modest valuation effects, such as the Canadian dollar. To the extent that the variable has predictive power for bilateral rates,

the Gourinchas-Rey model may have broader implications for empirical exchange rate modeling.

In several cases, the monetary model does significantly worse than the random walk, in a statistical sense. The result is consistent with Cheung et al. (2005a,b), indicating that the poor performance of the monetary models is not simply attributable to the extra noise introduced by the estimation procedure, as in Clark and West (forthcoming).

Overall, using appropriate test statistics yields the interesting finding that in 4 out of 42 cases significantly worse performance is recorded by the model-based prediction, while in 6 cases, significantly better performance is obtained (using a 10% marginal significance level). The latter ratio is more than would be implied by random occurrence.

## **6. Conclusion**

We find evidence in support of a new model of exchange rate modeling, motivated by intertemporal budget constraints, and contrast it with the results based upon interest rate parity and the sticky price monetary model. The evidence suggests that one can forecast changes in the spot rate using Gourinchas and Rey's *nx* measure and that interest rate parity outperforms the random walk at long horizons. The monetary model does not do well against the random walk, even though we rely upon a procedure for testing for statistical significance that has better size characteristics than the Diebold-Mariano test.

Finally, we leveled the approaches against a new currency, the euro. In this regard, we rely upon both actual data post-EMU and synthetic data that begin earlier.

In sum, we conclude:

- We cannot identify a model that reliably outperforms a random walk model, despite the use of an improved test.
- On the other hand, this better sized test forwarded by Clark and West indicates that the out-of-sample performance of structural models is not as poor as has been suggested by earlier statistical tests.
- The euro/dollar exchange rate – both its synthetic actual version – shares many of the same attributes that the deutschemark/dollar rate exhibited in terms of predictability and relevant determinants.
- The model that relies upon a log-linearized version of net exports variable does not do badly compared to the other models. However, perhaps for reasons related to data and sample differences, we are unable to exactly replicate the out-of-sample performance documented by Gourinchas and Rey.

The findings suggest immediate extensions and corrections. First, we hope to extend our estimates to multilateral effective exchange rates, although such an exercise is not straightforward because the nature of the effective exchange rates differs between models. Second, we intend to check the external imbalances approach using alternative net international investment position data. Third, we plan to extend the analysis to the Japanese yen/dollar rate. Finally, instead of relying upon interest rate parity, we can appeal to the entire term structure of interest differentials, as suggested by Clarida et al. (2001).

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## Appendix 1: Data

Unless otherwise stated, we use seasonally-adjusted quarterly data from the *IMF International Financial Statistics* ranging from the second quarter of 1973 to the last quarter of 2004. The exchange rate data are end of period exchange rates. The output data are measured in constant 2000 prices. The consumer price indexes also use 2000 as base year. Inflation rates are calculated as 4-quarter log differences of the CPI.

Canadian M1 and UK M4 are drawn from *IFS*. The US M1 is drawn from the Federal Reserve Bank of St. Louis's FRED II system. The euro area M1 is drawn from the ECB's Area Wide Macroeconomic Model (AWM) described in Fagan et al. (2001), located on the Euro Area Business Cycle Network website (<http://www.eabcn.org/data/awm/index.htm>).

The overnight interest rates are from the respective central banks. The three-month, annual and five-year interest rates are end-of-period constant maturity interest rates from the national central banks. For Canada and the United Kingdom, we extend the interest rate time series using data from IMF country desks. See Chinn and Meredith (2004) for details. We use German interest rate data from the Bundesbank to extend the Euro Area interest rates earlier.

The annual foreign asset and liability data are from the BEA website (<http://www.bea.gov/bea/di/home/iip.htm>). The quarterly data are interpolated by cumulating financial account flows from *IFS* and forcing the cumulative sum to equal the year end value from the BEA. The quarterly positions grow at the rate given by the financial account data in *IFS*, subject to the constraint that the year end value equals the official BEA data.

To construct the measure of external imbalance, we backed out the weights implied by the point estimates on page 13 of Gourinchas and Rey using the estimates from the DOLS regressions on page 12. The procedure assumes that the weights are constant across subsamples. The problem reduces to solving 3 equations in 3 unknowns, where the unknowns were the weights normalized on  $\mu_x$ . The weights are  $\mu_x / (\mu_x + 1) = 1.09639$ ;  $\mu_1 / (\mu_x + 1) = .72308$ ; and  $1 / (\mu_x + 1) = -0.092949$ .

Table 1: DOLS Estimates of the Monetary Model

	EUR	GBP	CAD
Money	-0.108 (0.297)	-0.192 (0.198)	0.112** (0.053)
Output	-0.996 (0.957)	-2.384 (1.597)	1.723*** (0.472)
Interest rates	-1.974 (1.325)	-0.634 (1.536)	-0.684 (0.805)
Inflation	8.102** (4.055)	1.196 (1.424)	2.547*** (1.193)
Adj. R-sq.	0.59	0.20	0.61
Sample	81Q4-04Q4	75Q4-04Q4	75Q2-04Q4
T	93	117	119

Notes: Point estimates from DOLS(2,2). Newey-West HAC standard errors in parentheses. \*(\*\*)(\*\*\*) indicates statistical significance at the 10% (5%) (1%) level.

Table 2: Interest Rate Parity Regressions

	EUR	EUR	GBP	CAD
3-month	-7.66*** (3.13)	-1.00 (1.65)	-1.58*** (0.99)	-0.84*** (0.47)
Adj. R-sq.	0.13	-0.01	0.02	0.01
Sample	99Q2-04Q4	90Q4-04Q4	75Q2-04Q4	75Q1-04Q4
T	23	104	119	120
1-year	...	...	-0.32 (0.82)	-0.57 (0.55)
Adj. R-sq.	...	...	-0.01	0.01
Sample	...	...	76Q2-04Q4	75Q1-04Q4
T	...	...	115	120
5-year	...	1.57 (0.38)	0.51 (0.37)	1.03 (0.31)
Adj. R-sq.	...	0.36	0.03	0.05
Sample	...	90Q1-04Q4	80Q1-04Q4	80Q1-04Q4
T	...	60	100	100

Notes: Point estimates from OLS. Newey-West HAC standard errors in parentheses. \*(\*\*)(\*\*\*) indicates statistical significance at the 10% (5%) (1%) level.

Table 3: Measure of External Imbalance

	EUR	GBP	CAD
3-month	-0.44*** (0.14)	-0.25* (0.13)	-0.14*** (0.07)
Adj. R-sq.	0.07	0.02	0.03
Sample	82Q1-04Q4	82Q1-04Q4	82Q1-04Q4
T	92	92	92
1-year	-0.42*** (0.10)	-0.30*** (0.08)	-0.16*** (0.05)
Adj. R-sq.	0.28	0.18	0.18
Sample	82Q1-04Q4	82Q1-04Q4	82Q1-04Q4
T	92	92	92
5-year	-0.20*** (0.05)	-0.14*** (0.04)	-0.08*** (0.02)
Adj. R-sq.	0.37	0.30	0.33
Sample	83Q4-04Q4	82Q1-04Q4	82Q1-04Q4
T	85	92	92

Notes: Point estimates from DOLS(2,2). Newey-West HAC standard errors in parentheses. \*(\*\*)(\*\*\*) indicates statistical significance at the 10% (5%) (1%) level.

Table 4: MSPE difference and Z-score for Clark-West Test

		(1)	(2)	(3)	(4)	(5)	(6)
		1987Q2-2004Q4			1999Q1-2004Q4		
	Horizon	IRP	S-P	NXA	IRP	S-P	NXA
Panel A: CAD/USD							
	1	-0.01 -0.35	-0.06 -0.70	-0.03 -0.49	-0.01 -0.37	-0.13 -1.62	0.10 2.21
	4	-0.08 -0.94	-0.12 -5.14	-0.06 -0.54	-0.14 -1.05	-0.09 -2.00	0.14 1.38
	20	0.50 0.30	0.08 0.09	0.06 0.22	0.27 3.22	... ...	... ...
Panel B: GBP/USD							
	1	-0.01 -0.11	0.09 -0.30	0.16 0.32	-0.02 -0.44	-0.13 -0.69	0.27 2.59
	4	-0.10 -0.38	0.22 3.42	-0.16 -0.19	-0.22 -1.23	-0.10 -2.12	0.33 1.09
	20	0.01 0.04	0.01 0.15	0.02 0.04	0.60 1.31	... ...	... ...
Panel C: EUR/USD							
	1	-0.01 -0.10	... ...	0.17 0.29	-0.02 -0.55	-0.73 -1.42	0.83 1.74
	4	... ...	... ...	-0.41 -0.45	... ...	-1.02 -6.52	0.70 1.22
	20	0.54 0.06	... ...	... ...	2.74 2.23	... ...	... ...

Notes: IRP: Interest parity model; S-P: Sticky price monetary model; NXA: Transformed net exports model. The first entry is the Clark-West test statistic times 100. The entry underneath is the z-score for the null hypothesis of no difference between the random walk and the linear alternative. For the interest rate parity forecasts, the z-score is the Diebold-Mariano (1995) test statistic. In some cases, 20-quarter ahead forecasts omitted because the Clark-West procedure requires the forecast sample be greater than two times the forecast horizon.

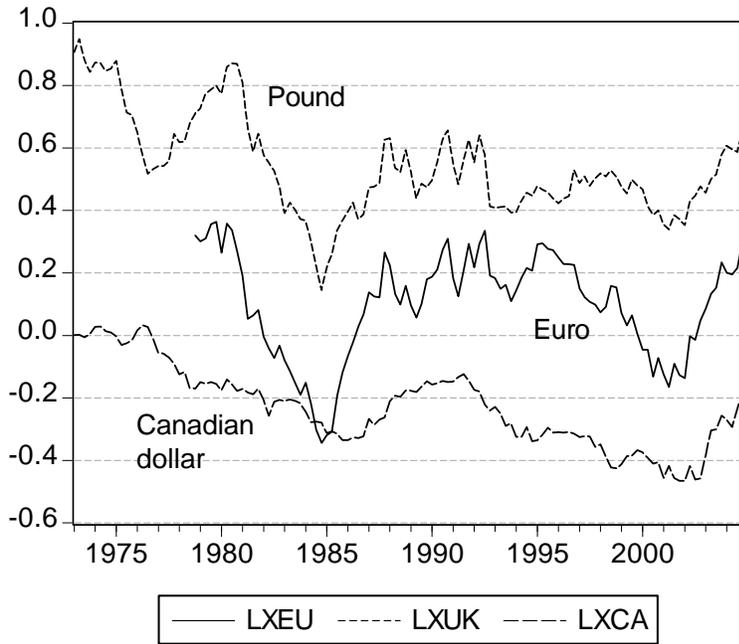


Figure 1: Synthetic and actual euro, UK pound and Canadian dollar exchange rates, end of quarter, in logs.

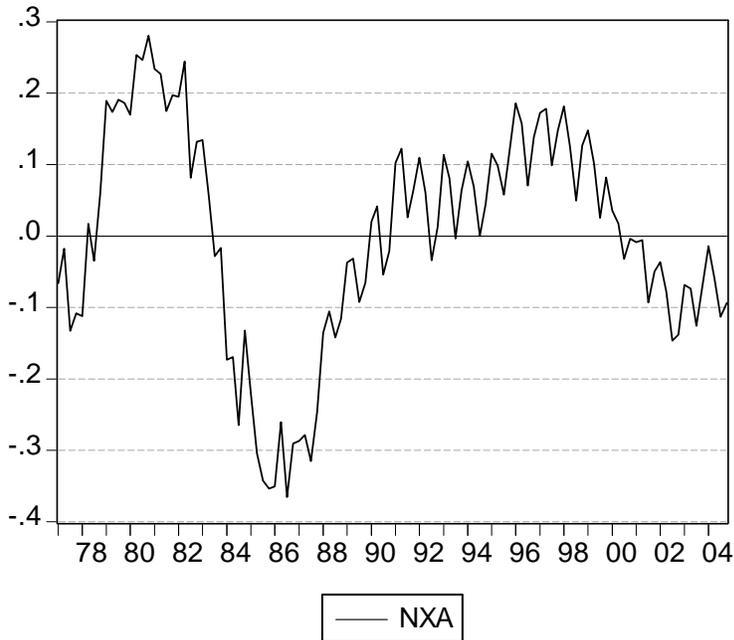


Figure 2: The Measure of External Imbalance.