Nominal Exchange Rates and Monetary Fundamentals: Evidence from a Seventeen Country Panel

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Abstract

We study the long-run relationship between nominal exchange rates and monetary fundamentals in a quarterly panel of 17 countries extending from 1973.1 to 1995.4. Our analysis is centered on two issues. First, we test whether exchange rates are cointegrated with long-run determinants predicted by economic theory. These results generally support the hypothesis of cointegration. The second issue is to re-examine the ability for monetary fundamentals to forecast future exchange rate returns. Panel regression estimates and forecasts confirm that this forecasting power is significant.

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Introduction

Recent studies have found that monetary fundamentals forecast nominal exchange rate returns (percent changes in the exchange rate) [see Mark (1995), Chinn and Meese (1995), Chen and Mark (1996), MacDonald and Taylor (1993), and MacDonald and Marsh (1997)]. Much of the evidence centers around significant slope coefficient estimates in regressions of future exchange rate returns on the deviation of the log exchange rate from monetary fundamentals and upon out-of-sample forecasts from these regressions which have been found to dominate the random walk in prediction accuracy. As was found in studies of stock returns [e.g., Fama and French (1988), Campbell and Shiller (1988), Hodrick (1992)], there is a tendency for regression slope coefficients and R^2s to increase in magnitude as the return horizon is lengthened. Similarly, out-of-sample forecast accuracy of monetary fundamentals relative to the random walk tends to improve with prediction horizon.

The relatively short time-series available for these studies combined with the high degree of dependence across overlapping observations of long-horizon exchange rate returns have made statistical inference a thorny issue and the robustness of the link between monetary fundamentals and the nominal exchange rate has been called into question by several authors. For example, in updating Mark's (1995) data set and employing a less restrictive data generating process to build the parametric bootstrap distributions upon which to draw inference, Kilian (1997) finds less favorable evidence for exchange rate predictability. Similarly, Berkowitz and Giorgianni (1997) and Berben and Van Dijk (1998) take the view that because exchange rates and their monetary fundamentals do not cointegrate, the independent variable in these regressions are nonstationary so that standard hypothesis testing procedures produce misleading inferences and need to be modified.[1] Moreover, all three authors argue that long-horizon regressions offer no statistical power gains over short-horizon regressions.

This paper is motivated by these recent critiques. We aim to improve on the imprecise univariate estimates and forecasts by exploiting available cross-sectional information in a panel data set. We analyze quarterly observations that begin on 1973.1 and extending through 1995.4 for 17 countries. We are focused on two main

[1] The fragility of the results evidently is not entirely due to the choice of statistical design. Groen (1996) finds considerable deterioration in the accuracy of out-of-sample monetary fundamentals forecasts of US dollar prices of the yen, deutschemark, and Swiss franc simply by extending Mark's quarterly sample, which ends in 1992, through 1994. We speculate that the collapse of univariate forecasts for the yen and the deutschemark in the 1990s may not be unreasonable. The ongoing banking crisis in Japan and residual fiscal consequences from German re-unification introduce important transient nonmonetary factors into the pricing of exchange rates.

issues. The first of these is whether nominal exchange rates are cointegrated with monetary fundamentals. To research this problem, we extend the principles for unit root testing in panel data suggested by Im, Pesaran and Shin (IPS) (1997) to tests for cointegration.

The second issue that we examine concerns the forecasting power of monetary fundamentals in a panel version of the long-horizon regression of currency returns on the deviation of the log exchange rate from the monetary fundamentals. To apply the standard regression analysis of stationary observations we must assume cointegration between the exchange rate and the fundamentals. On the other hand, if cointegration is not assumed, Berben and Van Dijk (1998) show that the univariate long-horizon regression tests can be interpreted as tests for weak exogeneity of the fundamentals. Under the null hypothesis of no cointegration, they show that the OLS slope coefficient has a nondegenerate asymptotic distribution which is a mixture of the normal and a Dickey–Fuller type distribution and the test of the hypothesis that the slope coefficient is zero is a test of joint null hypothesis of no cointegration and no predictability. While we do not employ Berben and Van Dijk's asymptotic analysis, we draw on their argument to justify our prediction analysis to readers who are unconvinced by the cointegration tests.

We first examine predictability by studying the estimated slope in the panel short-horizon regression. We confine this analysis to a one-period forecast period to avoid the complications arising from serially correlated error terms that would be induced by overlapping long-horizon forecasting horizons and to reflect the growing consensus that long-horizon regressions offer no statistical power advantages over short-horizon regressions. Secondly, we conduct an out-of-sample forecast experiment using the panel regression over the period extending from 1983.1 through 1995.4. In this analysis we report forecast results at two alternative horizons—one and sixteen quarters.

As typically done in forecast evaluation, we compare our fundamentals predictions to those of the random walk, but we also compare the exchange rate return forecasts by monetary fundamentals to forecasts by purchasing power parity (PPP). A growing body of empirical research employing panel methods on post 1973 data concludes that PPP holds over the float.[2] With the re–emergence of PPP as a viable long–run equilibrium condition for nominal exchange rate determination, it is logical and useful to implement an examination of its predictive power relative to the monetary fundamentals.

All of the statistical hypothesis tests that we conduct on the relevant regression coefficients and test statistics are based on parametric bootstrap distributions built

[2]See, for example, Frankel and Rose (1996), Lothian (1997), MacDonald (1996), Papell (1996) and Wu (1996).

on an estimated data generating process. We use the parametric bootstrap so as to control for potential finite sample size distortion of asymptotic tests. Additionally, the parametric bootstrap method allows us to model cross-sectional dependence in the data which surely is present in the data but which typically violates the regularity conditions under which the available asymptotic theory is derived.

The remainder of the paper is organized as follows. The next section describes the form of the monetary fundamentals predicted by theory that we use in our empirical analysis. Section 2 describes our data set. The panel cointegration and panel prediction analyses conducted on U.S. dollar exchange rates is covered in sections 3 4 respectively. Section 5 briefly discusses results using deutschemark and yen exchange rates. Section 6 concludes.

1 Monetary Fundamentals and the Exchange Rate

Let s_{it} be the time-t log nominal exchange rate between country $i = 1, 2, \ldots, N$ and the 'numeraire' country, which we label as country '0.' The exchange rate is country 0's currency price of a unit of currency i so an increase in s_{it} means a depreciation in value of 0's money. Similarly, let the time-t log nominal money stock and log real income of country i be denoted by m_{it} and y_{it} respectively. We refer to the long-run equilibrium exchange rate as the 'monetary fundamental value,' which we write as

$$f_{it} = m_{it} - m_{0t} - \alpha_i y_{it} + \alpha_0 y_{0t}, \tag{1}$$

where the long-run neutrality of money has been imposed and (α_i, α_0) are fixed parameters. Eq.(1) is a generic representation of the long-run equilibrium exchange rate implied by modern theories of exchange rate determination. The common feature shared by alternative theories is that the long-run equilibrium exchange rate is governed by determinants of money market equilibrium at home and abroad. In the monetary models of Frenkel (1976) and Mussa (1976), for example, α can be interpreted as the income elasticity of money demand and is predicted to be positive. In Lucas's (1982) equilibrium model, the sign of α, which may depend on preference parameters, is unrestricted but its value is bounded from above by 1.[3] The Obstfeld and Rogoff (1995) model on the other hand, predicts that per capita consumption

[3]In the Lucas model, the log equilibrium nominal exchange rate is given by $m - m_0 + (y + \ln U_y) - (y_0 + U_{y_0})$, where U_y and U_{y_0} is the representative agent's marginal utility of consumption. It is easy to see that a shock that raises y can lower the marginal utility of consuming y sufficiently to imply a negative value of α. For example, let the period utility function $U = (Y_0^{1-\gamma_0})/(1-\gamma_0) + (Y^{1-\gamma})/(1-\gamma)$. Then $\alpha = 1-\gamma$ and $\alpha_0 = 1-\gamma_0$. When the utility function displays curvature in excess of the log function, this model predicts negative α coefficients.

enter in place of real income. The particular details—whether to include income, as in Lucas, or consumption levels as in Obstfeld and Rogoff—differ, but the general theoretical prediction is the same. Namely that the exchange rate is determined by monetary fundamentals. This is not a controversial proposition but establishing this principle with a satisfactory degree of statistical accuracy has been difficult.

Our analysis centers on two empirical questions. First, we examine whether the monetary fundamentals given in (1) are an attractor for the nominal exchange rate. We address this problem in section 3 by testing whether whether $\{f_{it}\}$ and $\{s_{it}\}$ are cointegrated using procedures developed for the analysis of panel data. Second, we examine ability for deviations of the exchange rate from its monetary fundamentals value,

$$x_{it} \equiv f_{it} - s_{it}, \tag{2}$$

to forecast future exchange rate returns in a panel regression.

2 The Data

Our data consists of quarterly time series observations from 1973.1 through 1995.4 for the following 17 countries: Austria (AUT), Australia (AUS), Belgium (BEL), Canada (CAN), Denmark (DEN), Finland (FIN), France (FRA), Germany (DEU), Great Britain (GBR), Greece (GCE), Italy (ITA), Japan (JPN), Korea (KOR), the Netherlands (NET), Spain (ESP), Switzerland (SWI), and the United States (USA). The composition of the sample was determined by data availability and by a requirement that the country's exchange rate experience was dominated by floating. Some of the countries in the sample did experience episodes of nominal exchange rate pegging but those periods were deemed to have been reasonably brief.

Nominal exchange rates are end-of-quarter observations from the IFS CD-ROM (line code AE). Quarterly GDP is unavailable for several countries in the sample so as a result, we used quarterly industrial production indices for all countries as a proxy for national income. The industrial production series are from the OECD main economic indicators. Our measure of money is from the IFS and is the sum of money (line code 34) plus quasi-money (line code 35) for all countries with the following exceptions: Money is M0 for Great Britain and France and M3 for Spain. Price levels, which we need for the comparison to PPP, are measured using the CPI from the IFS (line code 66). Neither money nor the CPIs are seasonally adjusted. To control for seasonality, we filter the money and price series by applying a one-sided moving average of the current observation and 3-lagged values.

3 Panel Cointegration

To test for the presence of a unit root in a panel of observations $\{x_{it}\}$ of length T over N individuals, IPS (1997) propose the following method. Let t_i be the augmented Dickey–Fuller (ADF) studentized coefficient—the 't-statistic' that one computes—to test for a unit root in series i. The IPS panel unit root test statistic (their 'T-BAR' statistic) is formed by taking the cross-sectional average of the individual ADF statistics, $\bar{t} = (1/N)\sum_{i=1}^{N} t_i$. The null hypothesis is, all N time series contain a unit root. The alternative hypothesis is, **not** all N time series contain a unit root. The $\bar{t}$ statistic is asymptotically normally distributed under the null. Assuming that the cross-sectional units are stochastically independent, IPS provide calculations of the asymptotic mean and variance of $\bar{t}$.[4]

We test for cointegration between $\{s_{it}\}$ and $\{f_{it}\}$ by applying the principles developed by IPS. When we specify fundamentals a priori—α_i and α_0 chosen by us—the tests are a straightforward application of the IPS procedure. When (α_i, α_0) are estimated, we implement a two-step test of cointegration in which the $\bar{t}$ statistic is constructed from ADF regressions on residuals from a panel cointegrating regression. We bootstrap $\bar{t}$s to account for finite sample bias, cross-sectional dependence in the observations and for the fact that these tests are conducted on estimated residuals. The econometric specifications underlying our cointegration tests are described in subsections 3.1 and 3.2. Subsection 3.3 reports the empirical results.

3.1 A Two-Step Panel Cointegration Test

In our two-step tests we first run a panel cointegration regression and then apply the IPS test for a unit root in the regression residuals. The underlying econometric framework consists of a panel cointegrating regression, an ADF regression for the residuals, and a data generating process (DGP) that embodies the null hypothesis of no cointegration with which to build the parametric bootstrap distribution. We state the model here, then give an explanation of its components below.

$$e_{it} = c_i + \lambda_i t + \theta_t + \alpha_i y_{it} - \alpha_0 y_{0t} + z_{it}, \tag{3}$$

$$\Delta z_{it} = \delta_i z_{it-1} + \sum_{j=1}^{k_i} \gamma_{ij} \Delta z_{it-j} + u_{it}, \tag{4}$$

[4]Maddala and Wu (1996), develop a similar panel unit root test. They note that their method and the IPS procedure is a strategy originally suggested by Fisher (1932) for combining independent tests. See also, Pedroni (1998), who proposes a panel cointegration test based on nonparametric estimation of long-run autocovariance matrices.

$$
\begin{array}{rcl}
\Delta e_{it} & = & \mu_1^i + \sum_{j=1}^{k_i} a_{11,j}^i \Delta e_{it-j} + \sum_{j=1}^{k_i} a_{12,j}^i \Delta y_{it-j} + \sum_{j=1}^{k_i} a_{13,j}^i \Delta y_{0t-j} + \varepsilon_{1t}^i, \\
\Delta y_{it} & = & \mu_2^i + \sum_{j=1}^{k_i} a_{21,j}^i \Delta e_{it-j} + \sum_{j=1}^{k_i} a_{22,j}^i \Delta y_{it-j} + \sum_{j=1}^{k_i} a_{23,j}^i \Delta y_{0t-j} + \varepsilon_{2t}^i, \\
\Delta y_{0t} & = & \mu_0 + \sum_{j=1}^{k} a_{33,j} \Delta y_{0t-j} + \varepsilon_{0t}.
\end{array} \quad (5)
$$

Eq.(3) is the panel cointegration regression where we have imposed monetary neutrality by defining $e_{it} = s_{it} - (m_{it} - m_{0t})$ and regressing e_{it} on y_{it} and y_{0t}. We allow for the potential inclusion of three deterministic components in the regression. The first is an individual–specific constant c_i, that controls for the base year problem resulting from our use of industrial production indices to proxy for national income as well as for other non-time dependent heterogeneity in money market equilibrium across countries. The second is a deterministic trend $\lambda_i t$, that can be included to capture systematic changes in money market equilibrium across countries perhaps induced by differential rates of financial innovation. The third is a common time effect θ_t, that can be included to capture world–wide macroeconomic shocks. The common time effect also captures all numeraire country specific shocks, however, so α_0 is not identified when θ_t is included in the regression.

Eq.(4) is the ADF regression for the residuals from (3) and because the z_{it} are regression residuals, the ADF regression is run without constant or trend. The lag length k_i in the ADF regression is chosen by the general-to-specific sequential t-test method as suggested by Campbell and Perron (1991).[5]

Because our $\bar{t}$ is computed from estimated residuals, it will not have the asymptotic distribution derived by IPS. In addition, IPS's maintained assumption of independence across individuals is untenable in our study of bi-lateral exchange rates. Consequently, we cannot use their distributional results to perform cointegration tests. To test hypotheses, we generate a parametric bootstrap distribution for $\bar{t}$ with the system given in (5). The DGP is a restricted vector autoregression (VAR) for $(\Delta e_{it}, \Delta y_{it}, \Delta y_{0t})$ that embodies the null hypothesis of no cointegration between $\{s_{it}\}$ and $\{f_{it}\}$.[6] Eq.(5) is specified separately for each country. Ideally, we would like a DGP to jointly model the evolution of all of the variables across all 16 bilateral country pairs. This would imply an unrestricted VAR for 33 variables and would provide a proper accounting for the cross-sectional dependence across countries but estimating such a large system turns out not to be feasible. We should also point out that in our

[5]Starting with $k_i = 4$, estimate equation (4). If the absolute value of the t-ratio of the coefficient on the 4-th lag is less than 1.96, reset k_i to 3 and reestimate. The process is repeated until the t-ratio of the estimated coefficient with the longest lag exceeds 1.96.

[6]If $\{s_{it}\}$ and $\{f_{it}\}$ are cointegrated, the correct DGP would be given by a vector error correction representation.

VAR, the equation governing y_{0t} is restricted to a univariate autoregression. We do this to maintain uniqueness of the process governing y_{0t} since an unrestricted VAR would include lagged Δe_{it} and lagged Δy_{it} thus allowing the numeraire country's income process to vary across i.

The individual equations of the DGP are fitted by least squares with 4 lags. When we include linear trends in (3), we also include constants in (5). We account for dependence across cross-sectional units by estimating the joint error covariance matrix $\Sigma = E(\varepsilon_t \varepsilon_t')$ where $\varepsilon_t = (\varepsilon_{1t}^1, \ldots, \varepsilon_{1t}^N, \varepsilon_{2t}^1, \ldots, \varepsilon_{2t}^N, \varepsilon_{0t})$ from the OLS residuals.

We build the bootstrap distribution for $\bar{t}$ as follows. First, draw a sequence of length $T + 100$ innovation vectors is drawn according to $\tilde{\varepsilon}_t \sim N(0, \hat{\Sigma})$. Second, generate pseudo–observations $(\tilde{e}_{it}, \tilde{y}_{it}, \tilde{y}_{0t}), i = 1, \ldots, N, t = 1, \ldots, T + 100$ according to (5) using estimated values of the coefficients. Third, after dropping the first 100 pseudo-observations, run the cointegrating regression (3) on the pseudo–data and the ADF regression on the resulting regression residuals. This yields a realization of the IPS $\bar{t}$ statistic. Repeating these steps 2000 times yields the bootstrap distribution for $\bar{t}$.[7]

3.2 Cointegration Vector Fixed A Priori

If instead of estimating α_i and α_0 we set them to prespecified values, we can test for cointegration between $\{s_{it}\}$ and $\{f_{it}\}$ directly as suggested by IPS using

$$\Delta x_{it} = a_i + \lambda_i t + \theta_t + \delta_i x_{it-1} + \sum_{j=1}^{k_i} \gamma_{ij} \Delta x_{it-j} + u_{it} \tag{6}$$

and the DGP

$$\begin{aligned} \Delta s_{it} &= \mu_s^i + \varepsilon_{st}^i \\ \Delta x_{it} &= \mu_x^i + \textstyle\sum_{j=1}^{k_i} a_{21,j}^i \Delta s_{it-j} + \sum_{j=1}^{k_i} a_{22,j}^i \Delta x_{it-j} + \varepsilon_{xt}^i \end{aligned} \tag{7}$$

Here, the various deterministic components appear directly in the generalized ADF regression and the $\bar{t}$ statistic is built from the t_i computed from (6). The DGP, given in (7) is a restricted VAR in $(\Delta s_{it}, \Delta x_{it})$ in which the exchange rate follows a martingale and where $\{s_{it}\}$ and $\{f_{it}\}$ do not cointegrate.[8] Notice that the DGP allows for possible dependence of Δx_{it} on lagged Δs_{it}. Estimation of the DGP and buildup of the bootstrap distribution follows as described above in subsection 3.1.

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[7]An analogous procedure is followed to bootstrap all other statistics studied below and will not be repeated.

[8]It makes no difference whether the VAR is specified for $(\Delta s_{it}, \Delta x_{it})$ or $(\Delta s_{it}, \Delta f_{it})$ since $x_{it} = f_{it} - s_{it}$ and there is an independent relationship between only two of the three variables.

Table 1: Engle–Granger Cointegration Tests, ADF Unit Root tests

	Engle–Granger Tests				ADF Tests			
Co.	t_i	p-val	α_i	α_0	t_i	p-val	α_i	α_0
GBR	-2.578	0.664	0.588	-0.447	-2.409	0.454	1	0.1
AUT	-2.059	0.870	1.299	-0.524	-1.743	0.764	1	0.1
BEL	-1.262	0.982	1.279	-1.036	-0.866	0.942	1	0.1
DEN	-1.983	0.889	-0.823	-1.751	-1.647	0.790	1	0.1
FRA	-1.867	0.913	1.194	-0.720	-1.452	0.848	1	0.1
DEU	-1.245	0.983	1.531	-0.513	-0.996	0.929	1	0.1
NET	-2.096	0.860	2.474	-0.102	-1.909	0.700	1	0.1
CAN	-1.596	0.952	-1.933	-1.482	-2.569	0.380	1	0.1
JPN	-2.535	0.682	-0.927	-2.041	-1.636	0.795	1	0.1
FIN	-2.868	0.529	-0.252	-0.495	-2.767	0.294	1	0.1
GCE	-1.691	0.941	-0.289	-1.046	-1.352	0.876	1	0.1
ESP	-1.380	0.972	1.549	0.269	-1.392	0.870	1	0.1
AUS	-1.595	0.952	2.517	1.042	-1.418	0.861	1	0.1
ITA	-2.378	0.757	1.689	0.294	-2.130	0.593	1	0.1
SWI	-3.331	0.326	0.048	-1.292	-2.836	0.270	1	0.1
KOR	-2.112	0.854	-0.175	-1.313	-1.332	0.881	1	0.1

3.3 Cointegration Test Results

We use the U.S. as the numeraire country. To provide a benchmark for our panel results, table 1 shows the outcome of univariate Engle–Granger cointegration tests.[9] Here, it can be seen that the univariate tests provide little evidence for cointegration. The strongest evidence for cointegration is found for the US dollar–Swiss franc (p-value=0.33), followed by th US dollar–Finnish markka (p-value=0.53). The OLS estimates of α_i and α_0 display substantial heterogeneity across countries. Estimates of α_0 are predominately negative. The estimates of α_i are generally positive for European countries and are negative for Canada, Japan, and Korea. As discussed in section 1, Lucas's model bounds α_i and α_0 from above by 1 but does not restrict the sign of these coefficients.

[9]These tests were performed as described in section 3.1 except that the cointegrating regression is estimated individually for each exchange rate. We seem to get more sensible estimates when the cointegrating regression is run without trend we report only these results. We note, however, that the underlying strength of the evidence for cointegration from the univariate tests is not sensitive to inclusion of trend.

The last 4 columns of the table report unit root test results for fixed values of $\alpha_i = 1.0$ and $\alpha_0 = 0.1$. Our choice of these coefficient values is guided by the panel estimates which we discuss below. For now, we simply point out that the univariate unit root tests based upon fixed values of $(\alpha_i, \alpha_0) = (1.0, 0.1)$ only marginally strengthen the evidence against the unit root. Again, the strongest evidence for cointegration is for the dollar–Swiss franc (p-value=0.27) and the dollar–Finnish markka (p-value=0.30).

Panel A of table 2 shows the results of the two-step panel cointegration tests. When individual-specific fixed effects are included but no trends and no common time effect, we obtain a reasonably the large estimate $\hat{\alpha} = 0.86$ but a smaller estimate $\hat{\alpha}_0 = 0.14$ and little evidence of cointegration (p-value=0.45). Adding separate linear trends lowers the estimates and $\hat{\alpha}_0$ is now negative. The additional sampling variability introduced by the linear trends further weakens the evidence for cointegration (p-value=0.70). Presumably separate trends lead to overparameterizatoin of the model. An examination of the residuals revealed clear evidence of trend only for three countries–Great Britain, Japan, and Switzerland. To lighten the parameterization of the model, we included trends only for these countries with the result that evidence for cointegration becomes somewhat stronger (p-value=0.37).

To obtain reasonably strong evidence for cointegration, however, we need to control for common time effects. As mentioned above, estimates of α_0 are not available when θ_t is included in the regression. Point estimates of α appear to be robust to the inclusion of θ_t in the regression. When separate linear trends are included for all countries in addition to the common time effects, the hypothesis of no cointegration cannot be rejected at standard significance (p-value=0.23). Omitting the trends, or including them only for Great Britain, Japan, and Switzerland, allows the null hypothesis of no cointegration to be rejected at the 5 percent level (p-value=0.02).

Looking ahead to the next section, when we perform our out-of-sample forecasting experiment one option is to use the residual from eq.(3) z_{it} as the independent variable. As we will require recursive updates of z_{it}, we expect to achieve accurate forecasts if the estimates of (α, α_0) are reasonably stable over time. To examine the stability issue, figure 1 plots recursive panel estimates of (α, α_0). While estimates of α appear reasonably stable, the large dollar appreciation and subsequent depreciation of the 1980s and the relatively short time span of the data combine to make our estimates of α_0 display considerable sample dependence. To avoid these problems, we will pursue our analysis with a priori specified values of (α, α_0). To guide our choice of values, we refer back to the full-sample estimates in panel A of table 2 where we obtained relatively large estimates of α and relatively small but positive estimates of α_0. We therefore decided to fix $(\alpha, \alpha_0) = (1.0, 0.1)$.

In addition to the fragmentary empirical evidence, there are institutional reasons

Table 2: Panel IPS Cointegration Tests

Countries	α	α_0	Estimated	θ_t	λ_i	$\bar{t}$	p-value
			A. Two-Step Tests				
All	0.861	0.143	Yes	No	None	-1.665	0.450
All	0.161	-0.788	Yes	No	All	-2.097	0.701
All	0.802	0.117	Yes	No	GBR,JPN,SWI	-1.815	0.365
All	0.901	n.a.	Yes	Yes	None	-1.850	**0.042**
All	0.119	n.a.	Yes	Yes	All	-2.338	0.228
All	0.842	n.a.	Yes	Yes	GBR,JPN,SWI	-2.045	**0.019**
			B. Unit Root Tests with Pre-Specified (α, α_0)				
All	1.0	0.1	No	No	None	-1.763	0.203
All	1.0	0.1	No	No	All	-1.826	0.772
All	1.0	0.1	No	No	GBR,JPN,SWI	-1.888	0.215
All	1.0	n.a.	No	Yes	None	-1.907	**0.022**
All	1.0	n.a.	No	Yes	GBR,JPN,SWI	-2.133	**0.013**
No SWI	1.0	n.a.	No	Yes	GBR,JPN	-2.118	**0.038**
No SWI,FIN	1.0	n.a.	No	Yes	GBR,JPN	-2.138	**0.029**

that make the asymmetric treatment of the $\alpha-$coefficients—the relatively small value of α_0 in particular—sensible when the U.S. dollar serves as the numeraire currency. Because the U.S. dollar plays a the special role as a vehicle currency in international trade and finance, demand for U.S. money is comprised of at least three components of which the demand by U.S. residents is only one part. The other components of the demand for U.S. money arises from foreign Central Bank demand for foreign exchange reserves, and by foreign participants in international financial and trade markets.

Given these fixed values of the α–coefficients, we re-examine the evidence for cointegration. Panel B of table 2 shows the results of tests for a unit root with x_{it} constructed with $(\alpha, \alpha_0) = (1.0, 0.1)$. In comparison to the two-step tests, the evidence for cointegration largely unchanged when (α, α_0) are fixed. As in the earlier results, the common time-effect must be included in order to reject the null of no cointegration at the 5 percent level.

Finally recall that the Swiss franc and the Finnish markka displayed the most favorable evidence for cointegration in the univariate tests. We investigated the possibility that the panel test results are driven by the experience of these two series and conclude that they are not. Similarly, when both the Swiss franc and the Finnish markka (the currency with the second-most favorable evidence for cointegration) are

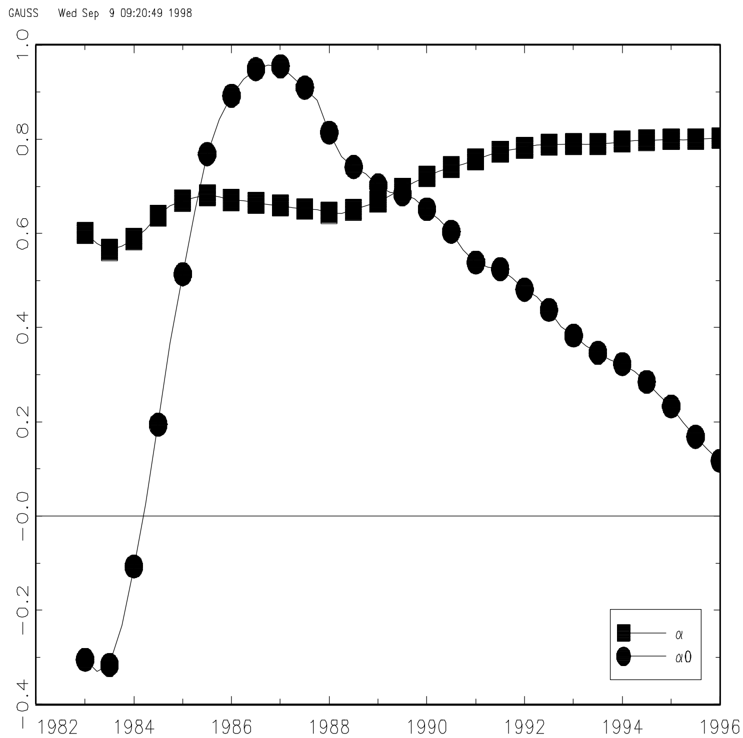

Figure 1: Recursive panel estimates of α and α_0.

dropped from the sample, we obtain a p-value of 0.022 and continue to reject the null of no cointegration.[10]

4 Prediction

Under the hypothesis of cointegration, the long-horizon regression is a regression of the stationary exchange rate return on stationary deviations from the monetary fundamental value x_{it}. In this case, the study of monetary fundamentals forecasts of exchange rate returns can proceed along the standard analysis of a panel regression with stationary variables.

If the hypothesis of cointegration is not maintained, the regression analysis of this section can be justified by appealing to Berben and Van Dijk (1998), who argue that under the null of no cointegration, testing for a zero slope coefficient in the long-

[10]Groen (1998) independently has rejected the null hypothesis of no cointegration between nominal exchange rates and monetary fundamentals in a panel of 14 OECD countries. His tests are based upon applying the Levin–Lin (1992) methodology on the residuals from a panel cointegrating regression.

horizon regression is equivalent to testing for weak exogeneity of the fundamentals. In the univariate setting assuming no cointegration between $\{s_{it}\}$ and $\{f_{it}\}$, they show that the OLS estimator in the univariate regression of Δs_{it} on x_{it} converges to its population value of zero and that its asymptotic distribution (and that of its asymptotic t-ratio) is a mixture of a Gaussian distribution and a Dickey-Fuller distribution which can be used to test the joint hypothesis of no cointegration and no predictability. Although asymptotic inference may in fact be feasible for the panel regression, we continue to conduct inference on forecast performance with the parametric bootstrap.

4.1 Predictive Regressions

In this section, we center our analysis on the panel regression,

$$s_{it+k} - s_{it} = a_i + \lambda_i t + \theta_{t+k} + \beta_k x_{it} + v_{it+k}. \tag{8}$$

In light of the instability displayed by recursive estimates of (α, α_0) and the relative insensitivity of the cointegration tests as to whether these coefficients are estimated or fixed at $(\alpha, \alpha_0) = (1.0, 0.1)$ as discussed in section 3.3, we construct the predictor x_{it} by setting the α–coefficients at these values. Eq.(8) is the panel analog of the long-horizon regression which possibly includes individual-specific (a_i) and time-specific (θ_t) fixed effects, as well as linear trends $(\lambda_i t)$. We focus on short-horizon estimates (k=1) first to avoid complications of residual serial dependence induced by forecasting over horizons exceeding the sampling interval of the data and second, to reflect the growing consensus that long horizons yield no power advantages over short horizons. Our first task here is to estimate $\beta_1 = \text{Cov}(\Delta s_{it+1}, x_{it})/\text{Var}(x_{it})$ and test the hypothesis $\beta_1 = 0$.

Following Berben and Van Dijk, eq.(8) is a legitimate regression whether or not $\{s_{it}\}$ and $\{f_{it}\}$ cointegrate. We therefore test the hypothesis $\beta_1 = 0$, under both assumptions. Under our first null hypothesis, the exchange rate evolves according to a martingale and is cointegrated with the monetary fundamentals. This null distribution is represented by the DGP,

$$\begin{aligned}
\Delta s_{it} &= \mu_s^i + \epsilon_{st}^i, \\
\Delta x_{it} &= \mu_x^i + \gamma_i x_{it-1} + \sum_{j=1}^{k_i} a_{21,j}^i \Delta s_{it-j} + \sum_{j=1}^{k_i} a_{22,j}^i \Delta x_{it-j} + \epsilon_{xt}^i.
\end{aligned} \tag{9}$$

We require $-2 < \gamma_i < 0$ so that eq.(9) is equivalent to a restricted VAR in $(\Delta s_{it}, x_{it})$ and which also has an equivalent a vector error-correction representation for $(\Delta s_{it}, \Delta f_{it})$

Table 3: Panel Short-Horizon Regression

β_1	p-value[1/]	p-value[2/]	t-ratio	p-value[1/]	p-value[2/]
0.055	**0.000**	**0.000**	6.818	**0.000**	**0.014**

Note: [1/] exchange rate follows a martingale and is cointegrated with fundamentals under the null. [2/] exchange rate follows a martingale but is not cointegrated with fundamentals under the null.

with cointegration vector $(1, -1)$. The estimation of the DGP and the buildup of the parametric bootstrap distribution follows the procedure described at teh end of subsection 3.1.

Our second null hypothesis is that the exchange rate evolves according to a martingale, but is not cointegrated with the monetary fundamentals. The DGP underlying this version of the null sets $\gamma_i = 0$ in eq.(9) (This DGP also was used previously and is stated in eq.(7)).

Table 3 shows the short-horizon panel regression results. Guided by our findings in the previous section, we control for the common time effect and include a linear trend for Great Britain, Japan, and Switzerland. Whether or not we assume cointegration under the null, the hypothesis of no predictability is rejected at standard significance levels.

4.2 Out-of-Sample Prediction

Out-of-sample forecasts are generated both at a short-horizon (k=1) and at a long-horizon (k=16). We begin by estimating eq.(8) on observations available through 1983.1. The k=1 regression is then used to forecast the 1-quarter ahead exchange rate return in 1983.2 and the k=16 regression to forecast the 16 quarter ahead exchange rate return through 1987.1. We then update the sample by one period by adding the observation for 1983.2 and repeat the procedure. Recursively updating in this fashion leaves us with 52 k=1 forecasts and 36 overlapping k=16 forecasts. Because we could not devise a satisfactory way to forecast future values of θ_t, we do not include the common time effect in generating the forecasts. This means that we cannot separate the contribution between the monetary fundamentals and the common time-effect for predictive performance. However, from the results of table 3, we can be reasonably confident that the role of the monetary fundamentals is significant.

To evaluate the monetary fundamentals, we compare the panel long-horizon regression forecasts against those of two alternative models. The first comparison is made against the standard benchmark forecast implied by the random walk model.[11] The other comparison that we make is against the predictions implied by PPP. Although many recent studies, using panel unit root tests on the log real exchange rate, have concluded that PPP has held over the float, few serious evaluations of the predictive performance of PPP have been undertaken.[12] These forecasts are generated by using deviations from PPP $[x_{it}^p = s_{it} - (p_{it} - p_{0t})]$ in the panel regression. The DGP for PPP forecast results is given in eqs.(7) and (9) with x_{it}^p in place of x_{it}. The PPP forecasts are similarly compared to the random walk with drift.

We evaluate forecast performance with Theil's U-statistic—the ratio of the root-mean-square prediction error (RMSPE) of the monetary fundamentals model (or alternatively the PPP model) to the RMSPE of the random walk. We avoided using other statistics of prediction evaluation, such as Diebold–Mariano (1995) because, as documented in Berben and Van Dijk, the difficulty in accurately estimating long-run variances often results in misleading inference. The null hypothesis is that the monetary fundamentals (or PPP) and the random walk provide equally accurate forecasts (U=1). The alternative hypothesis is that the monetary fundamentals (or PPP) is more accurate than the random walk (U<1). We also perform joint tests of the hypothesis of equal forecast accuracy by using joint test statistics formed alternatively by taking the mean value and the median value of the U-statistics. P-values are the proportion of the bootstrap distribution that lie below (to the left) of the U-statistic calculated from the data. As in the previous section, these p-values are constructed both under the null of cointegration and no cointegration.

The prediction results are displayed in table 4. At the 1-quarter horizon, monetary fundamentals point predictions dominate the random walk in RMSPE for 13 of 16 exchange rates (DEN, ITA, and ESP are the exceptions). The improvement in forecast accuracy is statistically significant at the 10 percent level for 5 exchange rates under the null of cointegration and for 4 exchange rates under the null of no cointegration. In addition, both of the joint tests can reject the hypothesis at the 10 percent level of equal prediction accuracy under the null of cointegration. The mean-value test statistic is also significant under the no cointegration null.

Whereas the 1-quarter random walk forecasts are more accurate than the monetary fundamentals for 3 of 16 exchange rates, the random walk dominates PPP in 8 cases. Looking at the last column of the table, it can be seen that monetary fun-

[11] We follow Kilian (1997) who argues that it is appropriate to employ the random walk with drift. We also evaluated the monetary fundamentals forecasts against the driftless random walk—a model which researchers often find performs better than the random walk with drift—but do not report those results to economize on space. For us, however, the results are robust to whether the drift is included.

[12] Chen and Mark (1996) and Lothian and Taylor (1996) are two exceptions.

Table 4: Out-of-Sample Forecasts

Country	U[a/]	p-value[1/]	p-value[2/]	U[b/]	p-value[1/]	p-value[2/]	U[c/]
A. One-quarter ahead forecasts							
AUS	**0.971**	0.123	**0.087**	**0.982**	**0.051**	0.192	**0.989**
AUT	**0.985**	0.135	0.260	1.000	0.274	0.527	**0.985**
BEL	**0.993**	0.109	0.421	1.001	0.316	0.590	**0.991**
CAN	**0.915**	**0.009**	**0.002**	**0.990**	0.141	0.356	**0.924**
DEN	1.000	0.193	0.510	1.006	0.471	0.673	**0.994**
GBR	**0.998**	0.256	0.307	**0.999**	0.328	0.548	**0.999**
FIN	**0.979**	**0.002**	0.173	**0.999**	0.302	0.545	**0.979**
FRA	**0.984**	**0.047**	0.244	1.007	0.621	0.686	**0.977**
GER	**0.993**	0.207	0.390	**0.997**	0.212	0.478	**0.996**
GCE	**0.989**	0.870	0.319	1.003	0.369	0.215	**0.987**
ITA	**0.971**	0.174	**0.082**	1.008	0.673	0.572	**0.964**
JPN	1.005	0.686	0.510	1.007	0.630	0.520	**0.998**
KOR	**0.759**	**0.001**	**0.001**	**0.991**	0.188	0.306	**0.766**
NET	**0.984**	0.101	0.236	**0.997**	0.219	0.491	**0.987**
ESP	1.001	0.485	0.566	1.016	0.686	0.682	**0.985**
SWI	**0.993**	0.289	0.193	**0.997**	0.218	0.456	**0.996**
Mean	**0.970**	**0.008**	**0.001**	1.000	0.165	0.211	**0.970**
Median	**0.989**	**0.042**	0.125	1.000	0.226	0.463	**0.987**
B.Sixteen-quarter ahead forecasts							
AUS	1.319	0.792	0.625	1.098	0.618	0.608	1.201
AUT	**0.398**	**0.006**	**0.007**	0.931	0.221	0.446	**0.428**
BEL	**0.961**	0.229	0.502	1.030	0.376	0.584	**0.933**
CAN	1.078	0.597	0.189	0.558	**0.009**	**0.060**	1.930
DEN	**0.606**	**0.017**	**0.084**	0.896	0.200	0.436	**0.677**
GBR	**0.532**	**0.023**	**0.006**	0.659	**0.028**	0.139	**0.807**
FIN	**0.553**	**0.011**	**0.050**	0.680	**0.029**	0.161	**0.813**
FRA	**0.395**	**0.001**	**0.005**	0.818	0.144	0.329	**0.483**
GER	**0.661**	**0.069**	0.134	0.988	0.294	0.523	**0.669**
GCE	1.293	0.937	0.766	1.712	0.937	0.469	**0.755**
ITA	**0.346**	**0.007**	**0.003**	0.927	0.306	0.380	**0.373**
JPN	**0.585**	0.137	**0.011**	1.116	0.642	0.527	**0.524**
KOR	**0.619**	0.184	**0.050**	0.556	**0.012**	0.046	1.114
NET	**0.380**	**0.004**	**0.006**	0.831	0.122	0.322	**0.457**
ESP	**0.540**	**0.063**	**0.043**	0.916	0.216	0.400	**0.589**
SWI	**0.606**	**0.082**	**0.014**	1.118	0.512	0.615	**0.542**
Mean	**0.679**	**0.003**	**0.001**	0.927	0.106	0.148	**0.768**
Median	**0.606**	**0.004**	**0.004**	0.927	0.151	0.317	**0.677**

Notes: [a/]: monetary fundamentals versus random walk with drift. [b/]: PPP fundamentals versus random walk with drift. [c/]: monetary fundamentals versus PPP. [1/] exchange rate follows a martingale and is cointegrated with fundamentals under the null. [2/] exchange rate follows a martingale but is not cointegrated with fundamentals under the null.

damentals point predictions dominate PPP predictions in RMPSE for all countries in the panel. Furthermore, under the null of cointegration, only one of the PPP U-statistics is significant at the 10 percent level while none are significant under the null of no cointegration. Similarly, neither of the joint tests reject the hypothesis of equal forecast accuracy at the 10 percent level.

At the 16-quarter horizon, monetary fundamentals forecasts again dominate the random walk in RMSPE for 13 of 16 exchange rates (AUS, CAN, and GCE are the exceptions). Under the null of cointegration, the hypothesis that the monetary fundamentals and the random walk provide equal forecast accuracy can be rejected for 7 exchange rates at the 5 percent level and for 10 exchange rates at the 10 percent level. Under the null of no cointegration, the hypothesis of equal forecast accuracy is rejected for 7 exchange rates at the 5 percent level and 11 exchange rates at the 10 percent level. Moreover, both joint test statistics are significant at the 5 percent level under either null hypothesis of cointegration.

PPP performance at 16 quarters continues to be inferior to monetary fundamentals. Monetary fundamentals forecasts dominate PPP forecasts in RMSPE for 13 of 16 exchange rates (AUS, CAN, and KOR being the exceptions). While PPP point predictions have lower RMSPE than the random walk for 11 of 16 exchange rates, the improvement in prediction accuracy is significant at the 10 percent level only for 4 exchange rates under the null of cointegration and for 2 exchange rates under the null of no cointegration.

5 Alternative Numeraire Countries

It has been reported in the literature that evidence for PPP is not invariant to choice of numeraire country [see Papell (199) and Wu (1997)] find evidence more favorable to PPP when Germany serves as the numeraire. To investigate whether numeraire dependence is present also in the monetary fundamentals–exchange rate link, we perform the econometric analysis on two alternative choices of numeraire: Germany and Japan. We give here a brief description of the major results.

5.1 Germany as Numeraire Country

Panel cointegration tests with Germany as the numeraire country are presented in table 5. As we found with dollar exchange rates, it is necessary to control for the common time effect in order to obtain reasonably strong evidence for cointegration. The panel estimates of α near 1 are on the high side, but drop substantially when trends are included in the panel cointegrating regression. Including trends for every

Table 5: Panel IPS Cointegration Tests with Germany as Numeraire Country

Countries	Estimated	α	α_0	θ_t	λ_i	$\bar{t}$	p-value
All	Yes	0.866	1.239	No	None	-1.670	0.165
All	Yes	0.482	0.397	No	All	-1.822	0.942
All	Yes	0.910	n.a.	Yes	None	-1.875	**0.032**
All	Yes	0.146	n.a.	Yes	All	-2.326	0.257
All	No	1.0	1.0	No	None	-1.969	0.662
All	No	1.0	1.0	No	All	-1.981	0.717
All	No	1.0	n.a.	Yes	None	-2.531	**0.043**
All	No	1.0	n.a.	Yes	All	-2.514	**0.088**
All	No	1.0	n.a.	Yes	11*	-2.365	**0.026**

Table 6: Panel Short-Horizon Regression

β_1	p-value$^{1/}$	p-value$^{2/}$	t-ratio	p-value$^{1/}$	p-value$^{2/}$
0.063	**0.001**	**0.031**	7.250	**0.000**	**0.054**

Notes: See notes to table 3.

country generates substantial sampling variability which weakens the evidence for cointegration.

Guided by the relatively large estimates of α, we run the unit root tests by fixing $(\alpha, \alpha_0) = (1.0, 1.0)$. Again, the evidence supports cointegration when the common time effect is controlled for. Including trend for all countries substantially weakens the evidence for cointegration. An examination of the residuals suggests that it is appropriate to include trend for 11 of the 16 countries.[13] Doing so for these 11 countries marginally strengthens the evidence for cointegration.

Table 6 reports estimates of the short-horizon panel regression. The null hypothesis of no predictability can be rejected at the 10 percent level under either assumption regarding cointegration.

The out-of-sample forecast results are shown in table 7. The monetary fundamentals forecasts perform less favorably for the deutschemark than for the U.S. dollar. At the 1-quarter horizon, monetary fundamental forecasts outperform the random walk for just 7 exchange rates and the improvement in accuracy is significant in 3 cases at the 10 percent level under the null of no cointegration. PPP forecasts perform

[13]They are, USA, GBR, BEL, DEN, FRA,CAN,JPN, GCE, AUS, ITA, and KOR.

Table 7: German Out-of-Sample Forecasts

Country	U^a	p-value[1/]	p-value[2/]	U^b	p-value[1/]	p-value[2/]	U^c
			A. One-quarter ahead forecasts				
AUS	**0.984**	0.279	0.107	**0.987**	**0.096**	0.254	**0.996**
AUT	1.555	0.972	0.668	1.022	0.680	0.779	1.522
BEL	1.110	0.990	0.837	**0.972**	0.039	0.079	1.142
CAN	1.021	0.753	0.542	1.000	0.225	0.443	1.021
DEN	1.142	0.985	0.868	1.043	0.881	0.832	1.095
GBR	**0.992**	0.201	0.094	**0.995**	0.210	0.330	**0.997**
FIN	**0.974**	0.484	0.033	**0.986**	**0.094**	0.224	**0.987**
FRA	**0.970**	0.474	0.180	1.002	0.307	0.394	**0.967**
GCE	**0.965**	0.963	**0.011**	**0.987**	0.104	0.085	**0.977**
ITA	1.022	0.915	0.678	1.023	0.796	0.516	**0.999**
JPN	**0.987**	0.202	0.125	1.013	0.687	0.693	**0.974**
KOR	**0.987**	0.334	0.131	**0.998**	0.222	0.383	**0.989**
NET	1.586	0.974	0.824	**0.942**	**0.002**	0.003	1.685
ESP	1.034	0.941	0.734	1.030	0.901	0.606	1.004
SWI	1.014	0.607	0.464	**0.971**	**0.021**	0.077	1.045
USA	1.023	0.720	0.606	**0.998**	0.201	0.425	1.026
Mean	1.085	0.994	0.607	**0.998**	0.182	0.230	1.089
Median	1.021	0.901	0.363	**0.998**	0.184	0.336	1.004
			B.Sixteen-quarter ahead forecasts				
AUS	**0.644**	0.145	**0.037**	**0.403**	**0.002**	**0.009**	1.599
AUT	10.070	1.000	0.965	4.779	0.999	0.999	2.108
BEL	1.528	0.864	0.715	1.086	0.509	0.494	1.407
CAN	**0.861**	0.295	0.175	0.549	**0.009**	0.044	1.567
DEN	2.063	0.963	0.874	1.685	0.874	0.819	1.224
GBR	1.164	0.698	0.478	1.464	0.855	0.736	**0.795**
FIN	1.326	0.682	0.207	1.284	0.618	0.656	1.033
FRA	**0.904**	0.717	0.478	1.298	0.691	0.611	**0.696**
GCE	**0.921**	0.952	0.263	**0.874**	0.191	0.112	1.055
ITA	1.184	0.858	0.642	1.625	0.909	0.690	**0.729**
JPN	**0.782**	0.200	0.108	1.632	0.944	0.879	**0.479**
KOR	**0.627**	0.138	**0.020**	1.224	0.681	0.628	**0.513**
NET	4.463	0.963	0.840	1.593	0.888	0.856	2.803
ESP	1.157	0.849	0.644	1.333	0.838	0.561	**0.868**
SWI	1.696	0.860	0.770	**0.420**	**0.002**	**0.009**	4.036
USA	**0.791**	0.154	0.125	**0.490**	**0.004**	**0.024**	1.614
Mean	1.886	0.988	0.881	1.359	0.779	0.675	1.408
Median	1.164	0.746	0.388	1.298	0.723	0.673	1.224

Notes: See notes to table 4.

slightly better, beating both the random walk and monetary fundamentals in terms of RMPSE for 8 of 16 exchange rates. PPP significantly (at the 10 percent level) outperforms the random walk for 5 exchange rates under the null of cointegration and for 4 exchange rates under the null of no cointegration, but neither of the joint test statistics are significant.

At the 16 quarter horizon, monetary fundamentals forecasts outperform the random walk in terms of RMPSE for 7 exchange rates, but only 2 of the U-statistics are significant at the 10 percent level under the null of no cointegration. These point predictions dominate PPP forecasts in only 6 cases. PPP forecasts outperform the random walk in terms of RMPSE for 5 exchange rates with 4 significant U-statistics under the null of no cointegration.

The evidence from the cointegration tests and the panel regression slope estimates is favorable to the monetary fundamentals, but the out-of-sample prediction evidence is spotty. Monetary fundamentals forecasts perform poorly for deutschmark exchange rates of European countries that participate with Germany in the Exchange Rate Mechanism of the European Monetary System. Point predictions by monetary fundamentals dominate the random walk for non ERM participants AUS, JPN, GCE, at the 1-quarter horizon and for AUS, CAN, GCE, JPN, KOR, and USA and at the 16 quarter horizon.

5.2 Japan as Numeraire Country

Panel cointegration tests for yen exchange rates are shown in table 8. The panel estimates of α_0 exceed 1 and are quite high but the estimates fall in size when separate trends are included. However, it is again necessary to control for the common time effect to obtain reasonably strong evidence for cointegration. When this is done, we estimate $\hat{\alpha} = 0.87$, and can reject the null of no cointegration at the 10 percent level (p-value=0.09).

In light of the relatively large estimates of α here, we run unit root tests by fixing $(\alpha, \alpha_0) = (1.0, 1.0)$. Again, the evidence supports cointegration when the common time effect is controlled for. Including trend for all countries only marginally weakens the evidence for cointegration. An examination of the residuals suggests, as it was for deutschemark exchange rates, that it is appropriate to include trend for 11 countries. When we do this, we obtain a p-value of 0.01 for $\bar{t}$.

Table 9 reports estimates for the short-horizon panel regression. The null hypothesis of cointegration and no predictability is easily rejected at very small significance levels. The null hypothesis of no cointegration and no predictability can also be rejected at the 10 percent level.

The out-of-sample prediction results for the yen are displayed in table 10. At

Table 8: Panel IPS Cointegration Tests with Japan as Numeraire Country

Countries	Estimated	α	α_0	θ_t	λ_i	$\bar{t}$	p-value
All	Yes	0.725	1.722	No	None	-2.384	**0.052**
All	Yes	0.010	-0.008	No	All	-2.301	0.724
All	Yes	0.873	n.a.	Yes	None	-1.887	**0.089**
All	Yes	0.200	n.a.	Yes	All	-2.370	**0.488**
All	No	1.0	1.0	No	None	-1.531	0.398
All	No	1.0	1.0	No	All	-2.808	**0.049**
All	No	1.0	n.a.	Yes	None	-1.934	**0.053**
All	No	1.0	n.a.	Yes	11*	-2.305	**0.005**

Table 9: Panel Short-Horizon Regression

β_1	p-value$^{1/}$	p-value$^{2/}$	t-ratio	p-value$^{1/}$	p-value$^{2/}$
0.040	**0.000**	**0.036**	5.848	**0.000**	**0.099**

p-value=0.000, no CI Notes: See notes to table 3.

the 1-quarter horizon, neither monetary fundamentals nor PPP offer impressive performance. Monetary fundamentals forecasts dominate the random walk in terms of RMPSE for only 4 exchange rates while PPP forecasts outperform the random walk for only 2 exchange rates. None of the individual nor the joint U-statistics are significant at the 10 percent level.

At the 16 quarter horizon, monetary fundamentals forecasts outperform the random walk in terms of RMPSE for 12 exchange rates. Under the null of cointegration, one of the U-statistics is significant (at the 10 percent level) whereas under the null of no cointegration, 8 U-statistics are significant, as are the two joint U-statistics. Moreover, the monetary fundamentals forecasts dominate PPP forecasts in terms of RMPSE for 10 of 16 yen exchange rates.

PPP forecasts outperform the random walk in terms of RMPSE for 9 exchange rates but only one of the individual U-statistics is significantly less than 1.

6 Conclusions

Univariate forecasts of exchange rate returns are imprecise. This paper aimed to exploit information in a small panel to are able to sharpen inference. Throughout the

Table 10: Yen Out-of-Sample Forecasts

Country	U^a	p-value$^{1/}$	p-value$^{2/}$	U^b	p-value$^{1/}$	p-value$^{2/}$	U^c
A. One-quarter ahead forecasts							
AUS	1.007	0.815	0.437	1.004	0.382	0.370	1.003
AUT	1.010	0.664	0.679	1.009	0.507	0.527	1.001
BEL	1.040	0.964	0.858	1.013	0.552	0.517	1.027
CAN	1.040	0.948	0.476	1.007	0.393	0.459	1.033
DEN	1.012	0.843	0.581	1.001	0.338	0.311	1.011
GBR	1.016	0.898	0.572	**0.996**	0.249	0.247	1.020
FIN	1.006	0.591	0.344	1.006	0.369	0.394	1.000
FRA	1.003	0.762	0.441	1.000	0.318	0.296	1.003
GER	**0.988**	0.223	0.134	1.014	0.566	0.609	**0.975**
GRE	1.015	0.987	0.605	1.013	0.602	0.162	1.002
ITA	**0.986**	0.775	0.105	**0.997**	0.293	0.205	0.989
KOR	1.041	0.958	0.816	1.001	0.418	0.330	1.040
NET	1.013	0.745	0.600	1.013	0.512	0.584	1.000
ESP	**0.988**	0.727	0.237	1.000	0.332	0.245	**0.988**
SWI	**0.974**	0.046	0.112	1.006	0.304	0.525	**0.968**
USA	1.018	0.839	0.556	1.003	0.342	0.432	1.014
Mean	1.010	0.907	0.257	1.005	0.323	0.297	1.005
Median	1.012	0.880	0.549	1.006	0.384	0.349	1.003
B.Sixteen-quarter ahead forecasts							
AUS	**0.696**	0.404	**0.070**	**0.920**	0.280	0.267	0.757
AUT	1.119	0.594	0.709	1.079	0.462	0.475	1.037
BEL	1.150	0.693	0.610	1.086	0.474	0.424	1.059
CAN	**0.571**	0.179	0.034	**0.775**	0.138	0.153	0.736
DEN	1.486	0.889	0.897	**0.756**	0.113	0.143	1.965
GBR	**0.739**	0.499	**0.074**	**0.767**	0.129	0.143	0.963
FIN	**0.609**	0.180	**0.095**	**0.786**	0.156	0.161	0.774
FRA	**0.887**	0.594	0.220	**0.772**	0.135	0.151	1.150
GER	**0.737**	0.171	**0.068**	1.331	0.698	0.674	0.553
GRE	**0.903**	0.933	0.243	1.124	0.545	0.196	0.803
ITA	**0.913**	0.788	0.258	**0.791**	0.141	0.134	1.154
KOR	1.376	0.877	0.782	**0.530**	**0.013**	**0.030**	2.596
NET	**0.702**	0.169	**0.043**	1.198	0.539	0.559	0.586
ESP	**0.460**	0.192	**0.025**	**0.759**	0.107	0.124	0.607
SWI	**0.683**	**0.059**	0.143	1.215	0.417	0.632	0.563
USA	**0.641**	0.181	**0.034**	**0.832**	0.185	0.227	0.771
Mean	**0.855**	0.453	**0.022**	**0.920**	0.191	0.213	1.005
Median	**0.739**	0.302	**0.005**	**0.832**	0.131	0.174	0.803

paper, statistical inferences were drawn using parametric bootstrap distributions of the relevant statistics.

The evidence for cointegration is strong when we control for common time effects. While the cointegration evidence is marginally weaker for deutschemark exchange rates than they are for either U.S. dollar or yen exchange rates, unlike some of the panel PPP literature, our results are not highly sensitive to the choice of the numeraire. The estimator of the common time effect is the cross-sectional average, so what is being tested is whether there are any series in the sample that are trending away from this time-varying mean. It bears mentioning that if the common time effect is in fact an I(1) random process, then the tests are misspecified. However, the correct way to sort out these issues is something that has not yet been addressed in the literature.

On the other hand, our analysis of the monetary fundamentals forecasts may not be all that sensitive to the properties of the underlying common time effect since valid statistical inference on the panel regression slope coefficient estimates can be drawn whether or not cointegration holds. The panel short-horizon slope estimates are highly significant under the null of cointegration, and are marginally significant, at the 5 percent level, under the null of no cointegration.

With regard to U.S. dollar exchange rates, one-period horizon point predictions dominated the random walk in root mean square error for 13 of 16 currencies. The fundamentals forecasts were significantly more accurate at the 10 percent level than the random walk for 5 exchange rates under the null of cointegration and for 4 exchange rates under the null of no cointegration. PPP forecasts at the one-period horizon on the other hand dominate the random in root mean square error only in 8 cases. Only one of the PPP forecasts is significant at the 10 percent level under the null of cointegration, and none are significant under the null of no cointegration. Monetary fundamentals forecasts of 16 quarter exchange rate returns are more dramatic. Monetary forecasts dominate the random walk in root mean square error in 13 cases for which forecasts for 10 currencies are significant at the 10 percent level. PPP forecasts on the other hand generate lower root mean square error in 11 cases but only 4 are significant at the 10 percent level.

We also considered two alternative numeraire currencies—the deutschemark and the yen. The cointegration and prediction evidence from the slope estimates appear firm. In out-of-sample analysis, long horizon point predictions from the monetary fundamentals dominate the random walk in 7 of 16 cases for the deutschemark and 12 of 16 cases for the yen but the gain in forecast accuracy is generally not statistically significant. While the evidence for the deutschemark and the yen is a bit fragmentary, the performance of PPP is roughly equivalent.

These results raise a pair of economic questions that are not solved by our analysis. First, while the slope on monetary fundamentals in the panel regression on deutschemark exchange rates is significant and is comparable to the dollar and yen results, out-of-sample monetary fundamentals forecasts perform badly against currencies of other European nations and especially so against countries that participate in the Exchange Rate Mechanism. Indeed, the worst performing forecasts are against the Dutch guilder—a currency that has been tightly tied to the deutschemark. A second puzzle is that the predictive content of monetary fundamentals are evidently superior to those of PPP in spite of the fact that PPP is the building block upon which the link between the exchange rate and monetary fundamentals are formed. Engel and Kim (1998), Canzoneri *et. al* (1996) report evidence that real exchange rates themselves contain relatively small, slow moving permanent components. A point of speculation for the poor PPP forecast performance may be the failure to account for embedded random walk dynamics.

References

Berben, R.B. and D.J. van Dijk (1998).'Does the Absence of Cointegration Explain the Typical Findings in Long Horizon Regressions?' Econometrics Institute, Erasmus University Rotterdam, Report 9814.

Berkowitz, Jeremy and Lorenzo Giorgianni (1997), 'Long-Horizon Exchange Rate Predictability?' International Monetary Fund working paper No. 97/6.

Campbell, John Y. and Pierre Perron, 1991, "Pitfalls and Opportunities: What Macroeconomists Should Know about Unit Roots," in O.J. Blanchard and S. Fisher, eds., *NBER Macroeconomics Annual*, Vol. 6, Cambridge, Mass., MIT Press.

Campbell, John Y. and Robert J. Shiller, 1988, 'Stock Prices, Earnings, and Expected Dividends,' *Journal of Finance*, July, 43, pp. 661-76.

Canzoneri, Matthew B., Robert E. Cumby, and Behzad Diba, (1996). 'Relative Labor Productivity and the Real Exchange Rate in the Long Run: Evidence for a Panel of OECD Countries,' *mimeo*, Georgetown University.

Chen, Jian and Nelson C. Mark, 'Alternative Long-Horizon Exchange Rate Predictors,' *International Journal of Finance and Economics*, 1,4 October 1996: 229–250.

Frenkel, Jacob A. "A Monetary Approach to the Exchange Rate: Doctrinal Aspects and Empirical Evidence." *Scandinavian Journal of Economics*, May 1976, 78, pp. 200–224.

Chinn, Menzie D. and Richard A. Meese, 1995, 'Banking on Currency Forecasts: How Predictable is Change in Money?' *Journal of International Economics*, 38, 161–178.

Diebold, Francis X. and Robert S. Mariano, 1995, 'Comparing Predictive Accuracy,' *Journal of Business and Economic Statistics*, 13, 253–263.

Engel, Charles and Chang-Jin Kim, (1998). 'The Long-Run U.S./U.K. Real Exchange Rate,' *Journal of Money, Credit, and Banking*, forthcoming.

Groen, Jan J.J. (1997). 'Long Horizon Predictability of Exchange Rates,' Tinbergen Institute Discussion Paper TI 97-095/2.

Groen, Jan J.J. (1998). 'The Monetary Exchange Rate Model as a Long–Run Phenomenon,' *mimeo*, Tinbergen Institute, Erasmus University, Rotterdam.

Fama, Eugene F. and Kennety R. French, 'Dividend Yields and Expected Stock Returns,' *Journal of Financial Economics*, October, 22, pp.3–25.

Fisher, R.A. (1932), *Statistical Methods for Research Workers.* Oliver and Boyd, Edinburgh 4th Ed.

Hodrick, R.J., 1992, Dividend Yields and Expected Stock Returns: Alternative Procedures for Inference and Measurement, *Review of Financial Studies* 5, 357–86.

Kyung So Im, M Hashem Pesaran and Yongcheol Shin (1997), 'Testing for Unit Roots in Heterogeneous Panels,' Discussion Paper, University of Cambridge.

Kilian, Lutz (1997). 'Exchange Rates and Monetary Fundamentals: What Do We Learn From Long-Horizon Regressions?' *mimeo*, University of Michigan.

Levin, Andrew and Chien-Fu Lin (1993). 'Unit Root Tests in Panel Data: New Results,' *mimeo*, Board of Governors of the Federal Reserve System.

Lothian, James R. (1997). 'Multi–Country Evidence on the Behavior of Purchasing Power Parity Under the Current Float,' *Journal of International Money and Finance*, 16, pp.19–35.

Lothian, James R. and Mark P. Taylor (1996), 'Real Exchange Rate Behavior: The Recent Float from the Perspective of the Past Two Centuries,' *Journal of Political Economy*, 104,

Lucas, R.E. Jr., (1982). 'Interest Rates and Currency Prices in a Two-Country World,' *Journal of Monetary Economics* 10, 335–359.

MacDonald, Ronald (1996). 'Panel Unit Root Tests and Real Exchange Rates,' *Economics Letters*, 50, pp.7-11.

MacDonald, Ronald and Ian Marsh (1997). 'On Fundamentals and Exchange Rates: A Casselian Perspective,' *Review of Economics and Statistics*, 79(4), November, pp. 655-64.

MacDonald, Ronald and Mark P. Taylor (1993). 'The Monetary Approach to the Exchange Rate: Rational Expectations, Long-Run Equilibrium and Forecasting', IMF *Staff Papers*, Vol. 40, No. 1, March.

Maddala, G.S. and Shaowen Wu (1996), 'A Comparative Study of Unit Root Tests wiht Panel Data and a New Simple Test,' mimeo, Ohio State University.

Mark, N.C. (1995). 'Exchange Rates and Fundamentals: Evidence on Long-Horizon Prediction,' *American Economic Review*, 85, 201–218.

Mussa, Michael. "The Exchange Rate, the Balance of Payments, and Monetary and Fiscal Policy under a Regime of Controlled Floating." *Scandinavian Journal of Economics* May 1976, 78, pp. 229–48.

Obstfeld, Maruice and Kenneth Rogoff, (1995), 'Exchange Rate Dynamics Redux,' *Journal of Political Economy*, 103 (June), pp. 624-60.

, Papell, David H. (1997). Searching for Stationarity: Purchasing Power Parity under the Current Float,' *Journal of International Economics*, 43, pp.313–332.

Pedroni, Peter (1997). 'Panel Cointegration: Asymptotic and Finite Sample Properties of Poolied Time Series Tests with an Application to the PPP Hypothesis,' *mimeo*, Indiana University.

Wu, Shaowen (1997). 'Purchasing Power Parity Under the Current Float: New Evidence from Panel Data Unit Root Tests,' *mimeo*, Ohio State University.

Wu, Yangru (1996). 'Are Real Exchange Rates Non-Stationary? Evidence from a Panel–Data Test,' *Journal of Money, Credit, and Banking*, 28, pp.54–63.