The Monetary Exchange Rate Model as a Long-Run Phenomenon*

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Abstract

Pure time series-based tests fail to find empirical support for monetary exchange rate models. In this paper we apply pooled time series estimation on a forward-looking monetary model, resulting in parameter estimates which are in compliance with the underlying theory. Based on a panel version of the Engle and Granger (1987) two-step procedure we find that the residuals of our pooled estimated model are stationary. This indicates that on a pooled time series level there is cointegration between the exchange rate and the macroeconomic fundamentals of this monetary model.

Keywords: monetary exchange rate models, nominal exchange rates, cointegration, panel data.
JEL classification: C23, F30, G15.

1 Introduction

The monetary model of the exchange rate is the standard instrument of analysis in international finance. In a way this is surprising as the empirical support for this economic model of exchange rate behavior is at the most doubtful. Although initially there was a claim of success for the monetary model, researchers quickly stumbled upon difficulties in finding an empirical fit for this model.

Frankel (1984) for example gives parameter estimates which do not concur with monetary exchange rate models, based on in-sample estimation over the period 1974-1981. The

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seminal paper by Meese and Rogoff (1983) shows that in out-of-sample forecasts monetary models do not beat the forecasting performance of naive random walk models. Recent papers like MacDonald and Taylor (1993) have tested the validity of monetary exchange rate models as long-run relationships for nominal exchange rates through cointegration techniques. Cointegration implies that a linear combination of non-stationary variables is stationary. MacDonald and Taylor find evidence for cointegration between the log of the exchange rate and the log of the fundamentals in a forward-looking monetary exchange rate model. Using the same techniques as MacDonald and Taylor, Sarantis (1994) concludes the opposite on a longer span of data. In general the papers on time series-based tests of monetary exchange rate models indicate that the inclusion of extra data points worsens the results. It is therefore difficult with pure time series to validate for the post-Bretton Woods era the appropriateness of monetary models as long-run phenomena.

In our paper we follow a different approach in that we look at pooled data sets of time series for at the most fourteen countries in the 1973-1994 post-Bretton Woods period. We use these panel data sets to ascertain the existence of long-run mean reversion in exchange rates based on a monetary exchange rate model under the current float. The failure of cointegration tests on time series of individual countries can be related to the availability of a short time span of data for the post-Bretton Woods floating period. Shiller and Perron (1985) in the context of unit root tests and Hakkio and Rush (1991) in the context of the Engle and Granger (1987) cointegration test, have shown that the power of these test procedures to reject the null in favor of the true alternative of stationarity or cointegration depends on the span of the sample. Extending the number of observations through an increase in the data frequency for the same time span does not improve the power of tests on unit roots or cointegration. An increase in the cross-section component of our sample increases the amount of long-run information in the data and should therefore improve the power of cointegration tests.

The panel data approach has become popular in international finance for testing the existence of purchasing power parity (PPP) in nominal exchange rates. Early examples of this approach are Hakkio (1984) and Koedijk and Schotman (1990), where both studies find evidence for a significant mean reversion component in real exchange rates by pooling data for four exchange rates. More recent studies are Abuaf and Jorion (1990), Frankel and Rose (1996), Oh (1996) and Papell (1997), where panel data sets are used with a cross-section dimension of at least ten exchange rates. The aforementioned studies give moderate evidence pro the existence of PPP in the recent post-Bretton Woods float. As PPP is a main building block of monetary exchange rate models, the panel data approach should also give more positive results for the long-run validity of these monetary models. A preliminary result of a possible success of pooling time series data for the monetary model can be found in Frankel (1984). Frankel pools the observations on a monetary exchange rate model for five bilateral dollar exchange rates and applies an ordinary least squares (OLS) regression on this stacked data set. Through this approach Frankel finds a moderate improvement in his parameters estimates compared with estimates for individual exchange rates. Our paper reports estimates of the static version of a simple monetary exchange rate model for our pooled data set and compares the point estimates with the theoretical values of the model parameters. As a next step we check if there is a stable long-run relationship in our panel data set between the exchange rate and the fundamentals of our structural model.

The set-up of this paper is as follows. In the second section we give an overview of the workhorse of our analysis, namely a flexible price, rational expectations version of the monetary exchange rate model. Section 3 contains pure time series based tests of our theoretical model for each individual country in our data set, followed by a pure cross-section regression approach. The fourth section deals with tests of the model for different panels of countries and their power to reject when the alternative hypothesis of cointegration is true. Conclusions and remarks can be found in section 5.

2 A monetary exchange rate model

Starting point for our model is the well known quantity relation, i.e. $M_tV_t = P_tY_t$ or in logarithms:

$$m_t + v_t = p_t + y_t, \tag{1}$$

where m_t, v_t, p_t and y_t are the logarithms in period t of the money supply, the velocity of circulation of money ("money velocity"), the price level and real income respectively. Suppose that PPP holds, that is $e_t = p_t - p_t^*$ where e_t, p_t and p_t^* are respectively the logarithms of the nominal exchange rate, the home country price level and the foreign price level.[1] Assume also that uncovered interest parity (UIP) holds: $i_t - i_t^* = E_t(e_{t+1}) - e_t$, where i, i^* are the home interest rate and the foreign interest rate respectively, while E_t is the conditional expectations operator. The two parity relations are combined with a log money velocity function which depends linearly on log real income and the interest rate:

$$v_t = \eta - \delta y_t + \omega i_t + \nu_t, \tag{2}$$

with η a constant, $\delta \geq 0$, $0 \leq \omega < 1$ and ν_t is a zero mean disturbance in the velocity function. One can see that (2) is the inverse of a standard money demand function, as money velocity is the inverse of money demand. Assuming that (1) and (2) holds at home and abroad with identical income elasticities δ and interest semi-elasticities ω, combining (1) and (2) with PPP and UIP results in the following expression for the log exchange rate e_t:

$$e_t = c + (m_t - m_t^*) - (1+\delta)(y_t - y_t^*) + \omega\left[E_t(e_{t+1}) - e_t\right] + \zeta_t. \tag{3}$$

[1] An asterisk indicates a foreign value of a variable, i.e. a US respectively a German value in sections 3 and 4.

In equation (3) $c = \eta - \eta^*$ and $\zeta_t = \nu_t - \nu_t^*$. Rearranging (3) and taking expectations, we get for $E_t(e_{t+1})$:

$$E_t(e_{t+1}) = \frac{c}{1+\omega} + \frac{1}{1+\omega}\left[E_t\left(m_{t+1} - m_{t+1}^*\right) - (1+\delta)\,E_t\left(y_{t+1} - y_{t+1}^*\right)\right] + \frac{1}{1+\omega}E_t(\zeta_{t+1}) + E_tE_{t+1}(e_{t+2}). \tag{4}$$

Using rational expectations, the law of iterated expectations and the zero mean property of ζ_t, recursive forward substitution based on (4) yields for (3):

$$e_t = c + \frac{1}{1+\omega}\sum_{i=0}^{\infty}\left(\frac{\omega}{1+\omega}\right)^i E_t(f_{t+i}) + \left(\frac{\zeta_t}{1+\omega}\right), \tag{5}$$

where:

$$f_t = (m_t - m_t^*) - (1+\delta)(y_t - y_t^*). \tag{6}$$

Here the fundamentals vector f_t consists in the country differentials of the log money supply and the log real income differential. Assume that f_t is a martingale (see section 3.1). This implies that $E_t(f_{t+1}) = \cdots = E_t(f_{t+k}) = f_t$. In that case equation (5) reduces to:

$$e_t = c + (m_t - m_t^*) - (1+\delta)(y_t - y_t^*) + \left(\frac{\zeta_t}{1+\omega}\right). \tag{7}$$

3 Time series tests and cross-country results

3.1 Individual time series tests

By now it is well known that time series tests of bilateral exchange rate models based on the monetary approach fail to find evidence pro specifications like (7). A major drawback of earlier studies, like Frankel (1984), is that they want to approximate short-run and long-run dynamics with a static regression as implied by (7). Next to that these studies disregard possible spurious regression effects due to non-stationarity in the elements of model 7. A model like (7) should more properly be interpreted as describing the long-run dynamics of the nominal exchange rate, as (7) is the outcome of rational expectations which should be valid on average. To test for mean reversion based on (7) we employ the cointegration technique of Johansen (1991) which is based on the vector autoregression (VAR) methodology to take into account short run dynamics. In this respect we follow MacDonald and Taylor (1993) and Sarantis (1994). The Johansen methodology uses estimated VAR's with error correction components to construct a number of likelihood ratio statistics equal to the number of endogenous variables K. These likelihood ratio statistics are used for testing the null hypothesis of at the most r cointegration vectors against the alternative hypothesis of $r+1$ cointegration vectors, with $r = 0, ..., (K-1)$.

We conduct cointegration tests for each i^{th} exchange rate on the following long-run model:

$$e_t = \beta_0 + \beta_1 (m_t - m_t^*) + \beta_2 (y_t - y_t^*) + \mu_t, \tag{8}$$

where according to the theory $\beta_1 = 1$ and $\beta_2 \leq -1$. The residuals μ_t must be stationary with a zero mean, as we can consider μ_t as the empirical proxy of $\left(\frac{\zeta_t}{1+\omega}\right)$ in (7). We use quarterly data over the period 1973:1-1994:4 for 14 bilateral exchange rates with respect to the United States (US) dollar, namely those of Australia, Austria, Canada, Finland, France, Germany, Italy, Japan, the Netherlands, Norway, Spain, Sweden, Switzerland and the United Kingdom (UK). Our monetary variables are based on the M1 definition, with the exception of Sweden for which we used relative M2. Relative real income is measured through Gross Domestic Product (Gross National Product in the case of Germany and Japan).[2] As noticed by Frenkel (1981) in individual time series and Papell (1997) in panel data sets, purchasing power parity seems to work better among European countries than for these European countries with respect to the United States. Given this result it would be interesting to see if our forward-looking monetary model also performs better with respect to an European numeraire. As most European countries pegged their exchange rates to the German Deutsche Mark (DM) during the post-Bretton Woods period, we repeat our analysis for (8) on a sample of fourteen DM-exchange rates. The fourteen DM rates and the corresponding macroeconomic fundamentals with respect to Germany are based on transformations of our original sample of fourteen US dollar rates and the corresponding macroeconomic fundamentals.[3]

In tables 1 and 2 we provide an overview of the empirical characterizations of the variables used in model (8). We characterize the data through the standard augmented Dickey-Fuller (ADF) unit root t-test to determine the order of integration. The ADF tests are applied with an intercept and intercept plus trend respectively. The number of lagged differences are selected by estimating the ADF regression without lags and then testing for serial correlation in the residuals with Lagrange Multiplier (LM) tests for one, four and eight lags. If we detect significant serial correlation we add a lagged first difference, estimate the regression model and utilize the LM tests again until there is no significant autocorrelation left in the residuals. In the case of the relative money supplies $(m_{it} - m_{it}^*)$ we use a different version of the ADF test, as $(m_{it} - m_{it}^*)$ is unadjusted for seasonal patterns. We follow Ghysels (1990) in that we replace the intercept with four seasonal dummies ($= d_s$):

$$\Delta x_t = \sum_{s=1}^{4} \chi_s d_s + \alpha x_{t-1} + \sum_{j=1}^{p} \phi_j \Delta x_{t-j} + \xi_t, \tag{9}$$

$$\Delta x_t = \sum_{s=1}^{4} \chi_s d_s + \Xi tr + \alpha x_{t-1} + \sum_{j=1}^{p} \phi_j \Delta x_{t-j} + \xi_t, \tag{10}$$

[2] For more details concerning the data used in this paper the reader is referred to Appendix A.

[3] See Appendix A.

Table 1: Unit root tests relative to the US, 1973:1-1994:4[a]

	e_t		$(m_t - m_t^*)$		$(y_t - y_t^*)$	
	const	trend	const	trend	const	trend
Australia	−1.529 (0)	−1.603 (0)	−0.108 (4)	−1.927 (4)	−1.805 (0)	−3.293 (1)
Austria	−1.079 (2)	−2.173 (3)	−1.223 (4)	−2.477 (4)	−2.094 (1)	−2.435 (1)
Canada	−1.020 (0)	−1.405 (0)	−1.192 (4)	−1.868 (4)	−3.804 (0)$^{\#}$	−2.649 (0)
Finland	−2.012 (5)	−2.042 (5)	−4.027 (4)$^{\#}$	−3.373 (4)	−1.322 (2)	−1.318 (2)
France	−1.468 (0)	−1.251 (0)	−2.055 (4)	−1.384 (4)	−1.736 (4)	−3.121 (2)
Germany	−1.507 (3)	−2.105 (3)	−1.757 (4)	−1.747 (4)	−1.617 (1)	−1.708 (1)
Italy	−1.664 (0)	−1.500 (0)	−2.427 (4)	−1.152 (4)	−2.355 (2)	−2.360 (2)
Japan	0.049 (0)	−2.492 (1)	−1.599 (4)	−2.866 (5)	−1.380 (0)	−1.360 (1)
Netherlands	−1.592 (3)	−2.038 (3)	−2.145 (4)	−4.586 (4)$^{\#}$	−1.731 (3)	−2.385 (3)
Norway	−1.676 (1)	−1.894 (3)	−1.222 (4)	−2.030 (4)	−3.177 (1)$^{\#}$	−3.063 (1)
Spain	−1.417 (1)	−1.379 (1)	−1.590 (4)	−3.232 (4)	−2.157 (1)	−2.231 (1)
Sweden	−1.524 (3)	−2.231 (3)	−1.471 (4)	−2.781 (4)	−1.268 (6)	−2.434 (0)
Switzerland	−1.817 (0)	−2.396 (0)	−0.125 (4)	−3.428 (4)	−2.438 (4)	−3.211 (4)
UK	−2.335 (3)	−2.551 (3)	−2.510 (5)	−2.927 (5)	−1.989 (0)	−2.611 (0)

[a] The number of lagged first differences are in parentheses. The term "const." ("trend") indicates an ADF regression with an intercept (intercept+trend). The symbol $^{\#}$ denotes significance of the ADF t-test at a 5 % level or lower.

Table 2: Unit root tests relative to Germany, 1973:1-1994:4[a]

	e_t		$(m_t - m_t^*)$		$(y_t - y_t^*)$	
	const.	trend	const.	trend	const.	trend
Australia	−1.472 (0)	−2.390 (0)	−0.542 (4)	−4.869 (4)$^{\#}$	−1.686 (0)	−1.531 (0)
Austria	−2.153 (8)	−2.311 (8)	−0.896 (4)	−2.370 (4)	−1.337 (0)	−2.419 (0)
Canada	−1.349 (0)	−2.509 (3)	−1.299 (4)	−1.560 (4)	−1.850 (1)	−1.483 (1)
Finland	−1.246 (0)	−2.372 (0)	−2.108 (4)	−1.016 (4)	−1.051 (2)	−1.125 (2)
France	−1.995 (1)	−1.021 (1)	−0.647 (4)	0.377 (4)	−0.025 (0)	−1.605 (0)
Italy	−1.663 (1)	−2.773 (0)	−1.700 (4)	0.255 (4)	−0.932 (1)	−1.238 (1)
Japan	−0.215 (5)	−2.843 (5)	0.305 (4)	−1.232 (4)	−1.618 (0)	−0.240 (0)
Netherlands	−1.475 (2)	−1.066 (2)	0.042 (4)	−1.233 (4)	−0.872 (1)	−2.749 (1)
Norway	−0.609 (0)	−2.593 (0)	−1.176 (4)	−1.110 (4)	−2.516 (1)	−1.806 (1)
Spain	−1.534 (0)	−1.945 (0)	−1.370 (4)	−1.314 (4)	−1.396 (1)	−1.892 (1)
Sweden	−0.640 (0)	−3.294 (1)	−1.452 (4)	−0.144 (4)	−0.055 (0)	−1.180 (0)
Switzerland	−1.943 (3)	−2.658 (3)	0.206 (4)	−1.525 (4)	−0.557 (0)	−1.347 (0)
UK	−1.937 (0)	−3.144 (0)	−0.944 (4)	−0.884 (4)	−1.390 (2)	−1.795 (2)
US	−1.507 (3)	−2.105 (3)	−1.757 (4)	−1.747 (4)	−1.617 (1)	−1.708 (1)

[a] See notes table 1.

Table 3: Individual cointegration tests, US dollar exchange rates, 1973:1-1994:4[a]

	Lags	LR(0)	LR(1)	LR(2)
Australia	5	14.30	3.70	0.34
Austria	4	25.13	7.96	2.03
Canada	4	27.24	5.22	0.86
Finland	4	35.63#	11.46	1.47
France	4	36.02#	17.69#	4.71#
Germany	5	26.81	9.64	2.09
Italy	4	27.00	12.59	4.24#
Japan	4	23.45	5.54	0.59
Netherlands	5	23.47	8.46	3.42
Norway	4	20.76	7.22	2.94
Spain	4	19.48	8.56	3.71
Sweden	4	32.93#	7.46	0.01
Switzerland	4	28.71	11.91	0.00
United Kingdom	4	20.48	7.22	1.00
5%		29.68	15.41	3.76
1%		35.65	20.04	6.65

[a] The term LR(r) denotes the likelihood ratio statistic for H_0: maximal r cointegrating vectors. The symbol # denotes significance of the likelihood ratio test at a 5% level or lower. The row "5%" ("10%") contains the appropriate 5% (10%) critical values.

where s is the index of the seasons ($s = 1, 2, 3, 4$) and tr is a time trend. In (9) and (10) we select $p = 4$ to correct for fourth order seasonal serial correlation and checked if based on LM(1), LM(4) and LM(8) tests this is appropriate. If necessary, we increased the number of lags beyond four until the LM tests indicated an absence of significant serial correlation in the residuals ξ_t in (9) and (10). Like Ghysels (1990) we use standard ADF critical values for inference in (9) and (10).[4] For all the ADF tests we use the MacKinnon (1991) critical values to test the null hypothesis of a unit root. The results in tables 1 and 2 indicate that in general, the variables that make up model (7) can be typified as I(1) processes. This confirms our martingale assumption for the fundamentals vector in the derivation of model (7). Hence, a cointegration frame work for testing mean reversion based on our monetary model (7) is appropriate.

In our cointegration tests we set the number of lags in our VAR models equal to four

[4] We checked this assumption through tabulation of critical values for (9) and (10) based on a Monte Carlo experiment with 10,000 iterations. The tabulated critical values were more or less equal to the appropriate MacKinnon (1991) critical values.

Table 4: Individual cointegration tests for DM exchange rates, 1973:1-1994:4[a]

	Lags	LR(0)	LR(1)	LR(2)
Australia	4	32.00#	7.49	0.67
Austria	5	22.08	6.91	0.51
Canada	4	17.70	4.30	1.33
Finland	4	38.25#	14.76	3.70
France	4	17.85	4.25	0.72
Italy	4	25.03	8.51	0.02
Japan	4	23.33	8.42	0.29
Netherlands	4	18.09	5.45	0.64
Norway	4	44.10#	17.86#	2.87
Spain	4	35.82#	17.13#	3.90#
Sweden	5	17.65	7.78	2.29
Switzerland	4	22.40	7.79	0.28
United Kingdom	4	24.64	8.68	0.25
United States	5	26.81	9.64	2.09
5%		29.68	15.41	3.76
1%		35.65	20.04	6.65

[a] See notes table 3.

to capture fourth order autocorrelation due to seasonality. After setting the lags equal to four and estimating the VAR's, we tested the residuals of the estimated VAR's on the absence of serial correlation with a Ljung-Box Q statistic with twelve lags and next to that tested through the Jarque-Bera statistic if there is normality in the residuals. If we detected significant serial correlation and non-normality we added a lag to our VAR's and run the residual based tests again. In our VAR's we included constants and three seasonal dummies, as our relative money supply variables are seasonally unadjusted.

Table 3 provides an overview of our cointegration tests per country for model (8) based on US dollar exchange rates. For a large majority of our bilateral exchange rates we cannot reject the null hypothesis of no cointegration. Only for Finland, France and Sweden are we able to find evidence for the existence of cointegration. Table 4 displays the results of cointegration tests on individual exchange rates based on our long-run model relative to Germany. Only in four out of our fourteen DM rates can we find evidence for cointegration on the basis of specification (8). All in all, our tests on time series for individual countries seems to indicate that our rational expectations model (7) inappropriately models the long-run behavior in nominal exchange rates.

3.2 Cross-country results

An alternative approach to analyze the long-run behavior between exchange rate movements and the movements in relative money supplies and real incomes within countries, is to exploit the cross-sectional aspects of the data. One advantage of this cross-section approach is that statistical inference of the estimated parameters is not influenced by issues like non-stationarity and cointegration. Another advantage of cross-section regressions is that the differences across countries result in a richer data set than in case of a time series approach. In our model long-run exchange rate movements are basically related to movements in excess money demand. Therefore, the ideal way to test our long-run model "...would be a comparison of long-term average behavior across economies with different monetary policies but similar in other respects" (Lucas 1980, p. 1006).

To analyze the cross-sectional aspects of our monetary exchange rate model we use the following regression model:

$$\Delta\bar{e}_i = \bar{\beta}_0 + \bar{\beta}_1 \Delta\tilde{m}_i + \bar{\beta}_2 \Delta\tilde{y}_i + u_i, \tag{11}$$

where $\Delta\bar{e}_i$ is the average quarterly exchange rate change for country i and $\Delta\tilde{m}_i$ respectively $\Delta\tilde{y}_i$ are the average quarterly changes in the relative money supply and relative real income with respect to the US or Germany for country i. Variable u_i is a zero mean disturbance and i is the country index, where $i = 1, ..., 14$. In table 5 one can find the regression estimates of (11) for both US dollar exchange rates and Deutsche Mark exchange rates, where the averages of quarterly changes are determined for the sample 1973:1 through 1994:4.

The estimation results in table 5 are for both cases quite favorable for our monetary exchange rate model, as the relative money parameter in both cases is not significantly

Table 5: Cross country regressions[a]

	US \$	DM
$\bar{\beta}_0$	-0.004 (0.001)	0.005 (0.001)
$\bar{\beta}_1$	1.093 (0.178)	1.024 (0.168)
$\bar{\beta}_2$	-2.229 (0.754)	-1.943 (0.725)
$t_{\bar{\beta}_1}$	0.522	0.143
$t_{\bar{\beta}_2}$	-0.307	0.079
R^2	0.775	0.771

[a] The OLS standard errors are in parentheses. Parameter estimates are denoted by $\bar{\beta}_0$, $\bar{\beta}_1$ and $\bar{\beta}_2$. The symbols $t_{\bar{\beta}_1}$ and $t_{\bar{\beta}_2}$ are the t-values for $\bar{\beta}_1 = 1$ and $\bar{\beta}_2 = -2$ respectively.

different from one and the relative real income coefficients are significantly negative. One remarkable feature of our cross country estimates is that for both US dollar rates and DM rates, the estimate for relative real income does not significantly differ from minus two. A parameter value $\bar{\beta}_2 = -2$ would suggest for model (7) that the income elasticity of real money demand equals one (as money velocity is the inverse of real money demand). Hence, our results indicate that there is a significant long-run relationship between the nominal exchange rate and our macroeconomic fundamentals with parameter values that are in accordance with the theoretical values in (7).

A graphical summary of our regression results can be found in figures 1 and 2. These figures represent the partial correlation between average exchange rate changes and average changes in relative money supplies and relative real incomes respectively. In constructing these diagrams we had to correct for the effect of the constant and the other variable not included in the scatter diagram, through partial regressions. For example, if we plot the partial correlation between the average changes of the exchange rate and relative money supplies, we first regressed both these variables on a constant and the average first differences of relative real income where the residuals of these regressions are the values of $\Delta\bar{e}_i$ and $\Delta\tilde{m}_i$ corrected for the effects of the intercept and $\Delta\tilde{y}_i$.[5] Obviously, the aforementioned procedure applies vice versa for the partial correlations between $\Delta\bar{e}_i$ and $\Delta\tilde{y}_i$.

In figures 1 and 2 we see that the scatter diagram confirm our regression estimates. The points in the scatter diagrams that represent the partial correlation between the average exchange rate depreciations and the average relative money growth rates, are clustered around a 45-degree line for both groups of exchange rates. This scatter pattern of exchange rate depreciation and relative money growth is in accordance with our theoretical model. One can see in all diagrams that the points are quite compactly clustered around the

[5] This is an application of the so-called Frisch-Waugh-Lovell theorem, see Davidson and MacKinnon (1993).

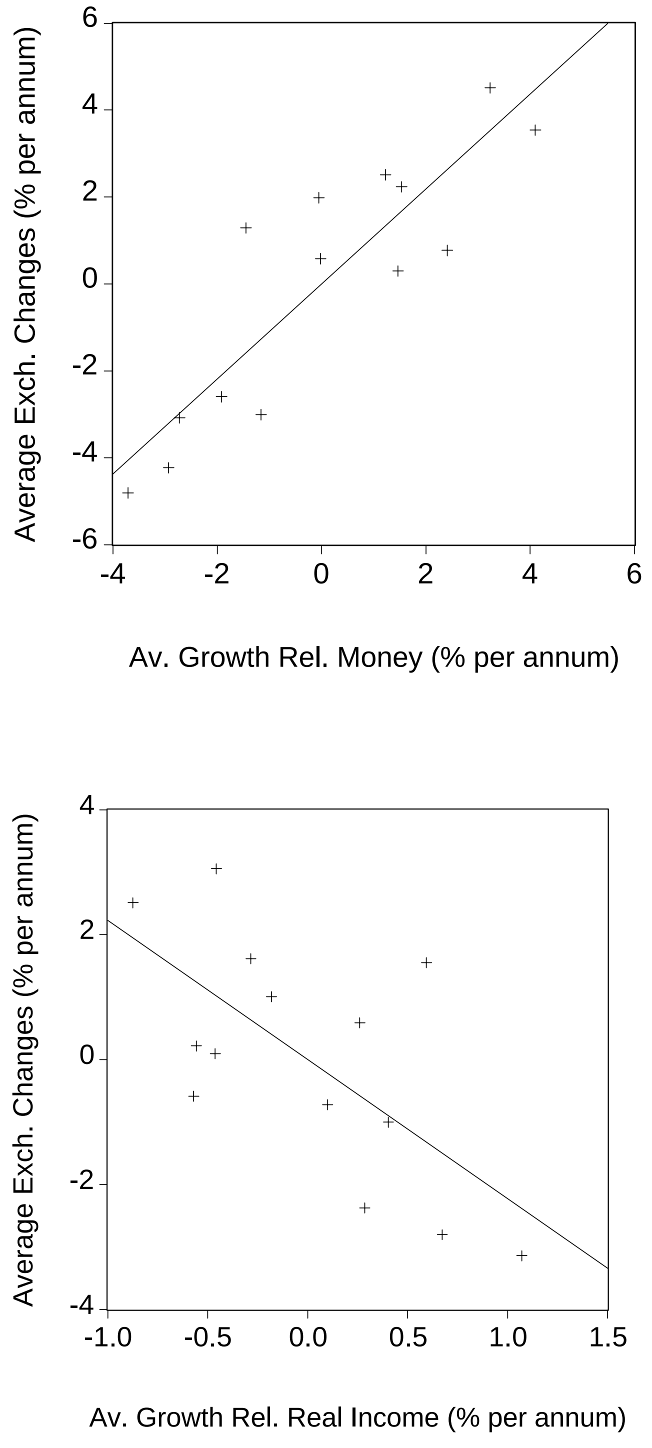

Figure 1: Partial cross-country correlations between the average quarterly changes of exchange rates and fundamentals, US dollar

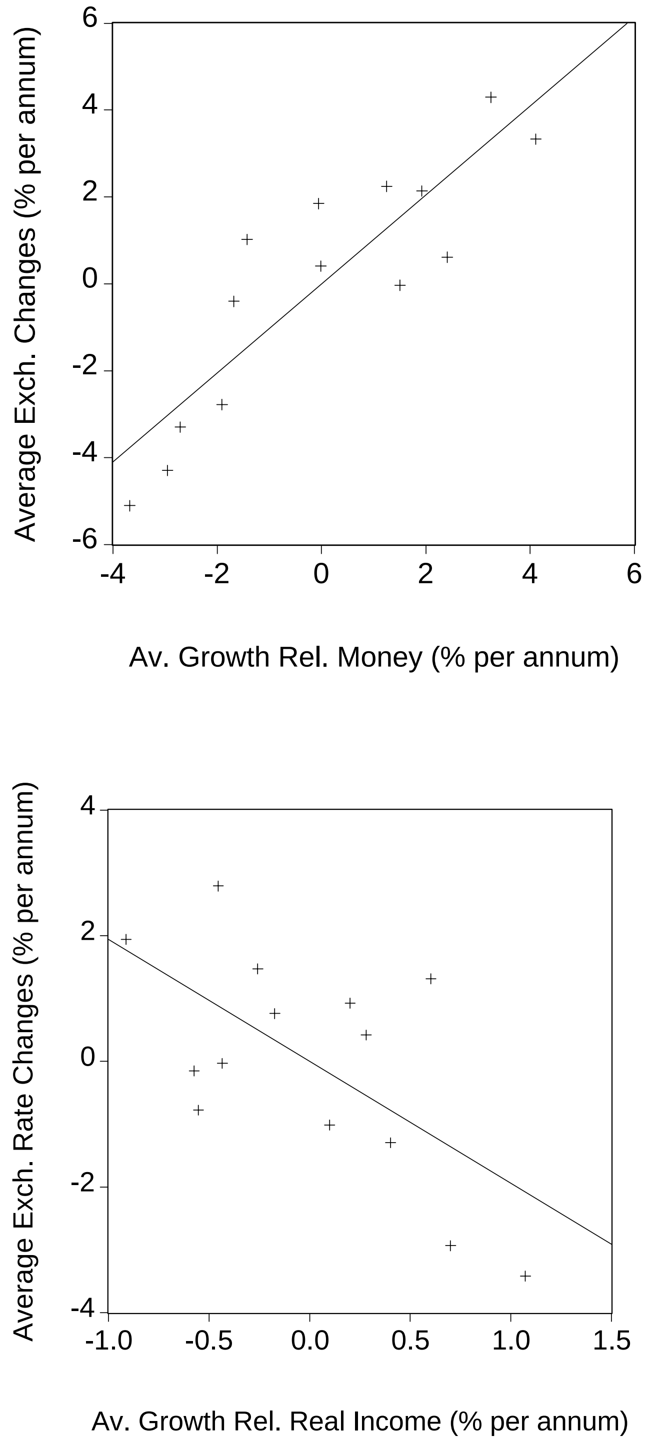

Figure 2: Partial cross-country correlations between the average quarterly changes of exchange rates and fundamentals, Deutsche Mark

partial regression lines, indicating that the assumption of homogeneous parameters across countries is not too far fetched.

4 A pooled time series approach

Time series-based tests for individual countries often reject a long-run forward-looking monetary exchange rate model like (7). A possible explanation is the lack of power in the available data, since the post-Bretton Woods floating period only spans 22 years in our data set. As numerous shocks have taken place during the mentioned period one has only available a handful of 'equilibrium' observations. An alternative approach would be to look at cross country regressions, as we did in subsection 3.2. Cross-section estimates indicate that our monetary model indeed has explanatory power in the long-run. A disadvantage of our cross country regressions is that in taking first differences and averaging over the sample, the information in time series is lost. In this section we want to ascertain if there is mean reversion over time based on our forward-looking monetary model (7). To this end we combine the dynamics in time series data with the long-run information in cross-sections into panel data sets of bilateral exchange rates and the corresponding set of macroeconomic variables. Four panel data sets are used for both US dollar rates and DM rates: all fourteen bilateral exchange rates of section 3, G10 rates, G7 rates and EMS rates. We use these panels to conduct tests on cointegration in a cointegrating regression based on a panel version of (8).

4.1 Methodology

In order to test for cointegration in our pooled data sets of exchange rate relationships, we apply like Pedroni (1995) the Engle and Granger (1987) two-step procedure on our panel data set. In the first step we run the following static regression on our pooled data set of exchange rates:

$$e_{it} = \beta_{0,i} + \sum_{s=2}^{4} \chi_{is} d_{is} + \beta_1 \left(m_{it} - m_{it}^{*} \right) + \beta_2 \left(y_{it} - y_{it}^{*} \right) + \mu_{it}, \tag{12}$$

where d_{is} are three seasonal dummies to correct for the effect of seasonality in $(m_{it} - m_{it}^{*})$, i is the indicator of the cross-section index and t is the time index. As in (8), β_1 is expected to be 1 and β_2 to be negative. Like Frankel (1984), we have constrainted the parameters of the relative money supplies and relative real incomes to be identical across the exchange rates. The intercept $\beta_{0,i}$ and the seasonal effects can vary cross-sectionally in (12).

As a second step we can check if the estimated residuals of (12) ($= \hat{\mu}_{it}$) are non-stationary. In order to do that we apply the Augmented Dickey Fuller (ADF) regression to $\hat{\mu}_{it}$:

$$\Delta \hat{\mu}_{it} = \alpha \hat{\mu}_{i,t-1} + \sum_{j=1}^{p} \phi_{ij} \Delta \hat{\mu}_{i,t-j} + \varepsilon_{it}. \tag{13}$$

In this regression $\Delta\hat{\mu}_{it} = \hat{\mu}_{it} - \hat{\mu}_{i,t-1}$ and p indicates the number of lagged first differences. Equation (13) tests the null hypothesis of non-stationarity in $\hat{\mu}_{it}$, i.e. the null hypothesis is that of non-cointegration. The alternative hypothesis is that of mean-reversion of $\hat{\mu}_{it}$ to zero or in other words: the alternative is that of cointegration. Therefore, the null hypothesis is tested through an one-sided t-test for $\alpha = 0$ against the alternative hypothesis $\alpha < 0$. As we have estimated the cointegrating vector on a pooled basis to get more powerful parameter estimates, we also have parameter α constrained to be cross-sectionally identical to get more powerful inference with respect to the null hypothesis of non-stationary residuals $\hat{\mu}_{it}$.

Based on a cointegrating regression with a common intercept, cross-sectional independence in the residuals and a common residual variance, Pedroni (1995) shows that our test with $p = 0$ has the following asymptotics as T and N go to infinity (T and N are the number of time series and cross-sections respectively and $\sigma(\alpha)$ is the standard error of α):

$$\begin{aligned} T\sqrt{N}\hat{\alpha} &\Rightarrow N(0,2), \\ \hat{\sigma}(\hat{\alpha}) &\Rightarrow \sigma(\alpha), \\ t_\alpha \left(= \frac{\hat{\alpha}}{\hat{\sigma}(\hat{\alpha})} \right) &\Rightarrow N(0,1). \end{aligned}$$

The asymptotics based on Pedroni's assumptions are the same as those of a panel unit root test without individual intercepts (Levin and Lin 1992). In contrast to Pedroni, we use in (12) individual intercepts, resulting in a mean shift in t_α. If we maintain Pedroni's assumptions of cross-section independence and a common variance in ε_{it}, our employed test has the same limit distribution as a panel unit root test with individual intercepts:

$$\sqrt{1.25}t_\alpha + \sqrt{1.875N} \Rightarrow N(0,1).$$

Hence, the use of individual intercepts induces a mean shift in the distribution of t_α.

We allow for serial correlation of the disturbances in $\hat{\mu}_{it}$ and assume this to be heterogeneous across the exchange rates. In contrast to Koedijk and Schotman (1990) and O'Connell (1998), our test results are not numeraire invariant due to the assumption of heterogeneous serial correlation and the results may therefore differ for US dollar rates and DM rates. Unlike Pedroni (1995) we apply a parametric correction for possible serial correlation, as Pedroni uses a non-parametric correction due to Phillips and Perron (1988). The number of lagged differences in (13) is assumed to be four ($p = 4$) in all panels to correct for fourth order seasonal autocorrelation due to seasonality in the relative money supplies. To correct for level effects of the seasonality in $(m_{it} - m_{it}^*)$ we included seasonal dummies in the cointegrating regression (12). For the US dollar exchange rates we found that in ten out of fourteen bilateral exchange rates at least the first two lagged first differences where significant, with a maximum of four. In the case of DM exchange rates the aforementioned phenomenon occurred for seven out of fourteen bilateral exchange rates, again with a maximum of four lagged first differences. Using LM serial correlation tests at one, four and eight lags, (13) with $p = 4$ seems to deal with all the serial correlation in the disturbances in (13) for each i^{th} exchange rate in all our panel data sets. The choice of $p = 4$ seems therefore to be justified.

With individual time series one estimates the ADF equation with OLS. In principle one could also do that with respect to the panel residual ADF in (13). A feature of exchange rates is that the relative changes in exchange rates are contemporaneous correlated across the countries, as can be seen in table C.1 in the Appendix. This is not surprising as our exchange rates have the same numeraire, in this case the US dollar. We can relate this cross-section correlation to the presence of the US money demand shock ν_t^* in the residuals $\hat{\mu}_{it}$ for each country i, as is assumed in our forward-looking model (7). From table C.2 one can see that the cross-section correlation in the exchange rate changes induces a comparable amount of cross-sectional dependence in the OLS residuals of (13), after estimating (12) on our fourteen US dollar exchange rates and set $p = 4$ in (13). Tables C.3 and C.4 contain comparable results for our DM exchange rates.

Not correcting for the aformentioned cross-section dependence in the ε_{it}'s of (13) severely affects the statistical properties of our panel cointegration test. As an illustration we conduct a Monte Carlo experiment to asses the size of cointegration tests if one does not take cross-section dependence into account. In our experiment we generate the non-stationary dependent variable through the following data genrating process (DGP):

$$y_{it} = \theta_i + x_{1,it} - 2x_{2,it} + \phi_{it},$$

where the individual intercepts θ_i's are generated on a one time basis for each panel through an uniform distribution $U\,[0, 10]$ and our explanatory variables $(x_{1,it},\ x_{2,it})$ are generated as independent random walks with standard normal innovations. The residuals ϕ_{it} are also assumed to be random walks with standard normal innovations, albeit that these innovations have a cross-section correlation equal to γ. We generate for each panel four sets of ϕ_{it}'s (and therefore four sets of y_{it}'s): with the cross-section correlation in the innovations of ϕ_{it} equal to $\gamma = 0$, 0.2, 0.5 and 0.9. For these four cases we apply the aforementioned two-step panel cointegration procedure (with $p = 0$) where we regress the different y_{it}'s on $x_{1,it}$ and $x_{2,it}$, and create four Monte Carlo distributions of t_α. The t_α's based on $\gamma = 0$ are used to derive the critical values, while the distributions of t_α with cross dependence are used to asses the size of these critical values. We conduct the size experiments for three panels which have more or less the same dimensions as in our empirical analysis: all three panels have 100 time series observations and respectively 5, 10 and 15 cross-sections. The size ratios based on 10,000 Monte Carlo experiments can be found in table 6.

Table 6 indicate that cross-section dependence in ε_{it} of (13) severely distorts the size of tests based on cross-section independence. It is clear that the larger the cross correlation the higher are the size ratios for the 1%, 5% and 10% critical values. Therefore, estimation of (13) should take into account the cross-section correlation in ε_{it} of (13) otherwise the derived critical values are inappropriate.

Pedroni (1995) uses common time dummies in a cointegration regression like (12) to correct for the aforementioned cross-section dependence. With common time dummies

Table 6: The size of panel cointegration tests with neglected cross-section dependence[a]

			True size given nominal size:		
N	T	γ	1%	5%	10%
5	100	0.2	0.014	0.059	0.113
5	100	0.5	0.031	0.090	0.156
5	100	0.9	0.130	0.231	0.314
10	100	0.2	0.015	0.067	0.124
10	100	0.5	0.059	0.150	0.223
10	100	0.9	0.241	0.352	0.414
15	100	0.2	0.018	0.069	0.123
15	100	0.5	0.091	0.187	0.261
15	100	0.9	0.325	0.416	0.468

[a] N is the number of cross-sections. T is the number of time series observations. The symbol γ indicates the measure of cross-section correlation in the innovations of residuals ϕ_{it} (see text). The size ratios are based on 10,000 trials of the Monte Carlo experiment described in the text.

the off-diagonal elements in the cross-sectional covariance matrix of ε_{it} in (13) decreases to an order $\frac{1}{N-1}$ of the diagonal elements (O'Connell 1998). Thus in panels with a moderate or small number of cross-sections, common time dummies do not eliminate cross-section dependence completely. As an alternative we estimate (13) with feasible generalized least squares (FGLS) instead of OLS, based on a transformation of the variables in the regression. In this transformation we postmultiply the left hand and right hand variables in (13) with a Cholesky decomposition of the inverse of the estimated cross-section covariance matrix $E\left(\hat{\varepsilon}_t\hat{\varepsilon}_t'\right)$, where $\hat{\varepsilon}_t = [\hat{\varepsilon}_{1t}, ..., \hat{\varepsilon}_{Nt}]$ and $\hat{\varepsilon}_{it}$ are the OLS residuals of (13). Estimation with FGLS does not only eliminate cross-sectional dependence in ε_{it} completely, it also yields the result that the FGLS residuals of (13) have a unit variance for every exchange rate i. Therefore, our FGLS t-ratio of α is more in line with the above mentioned asymptotic assumptions of Pedroni (1995).

Finite sample critical values for the FGLS t-value on α in (13) are tabulated by a parametric bootstrap procedure. As outlined in Appendix B, we assume that the fundamentals and the residuals of (12) are martingales. To generate these martingales we draw bootstrap samples of the innovations in the fundamentals and μ_{it} based on fitted AR models, where the innovations of μ_{it} have cross-section dependence. As a next step we use the generated series of the fundamentals and μ_{it} to construct artificial exchange rate data. The simulated samples of the exchange rates and the fundamentals are then used to apply our panel cointegration test procedure. We determine the appropriate finite sample distributions for the FGLS t_α through 20,000 replications of the above mentioned parametric bootstrap procedure.

4.2 Test results

Our panel data set is constructed by pooling the fourteen quarterly bilateral dollar-exchange rates and their corresponding fundamentals of section 3. We use the two-step procedure outlined in subsection 4.1 to test for cointegration in our forward-looking monetary model. As a first step we estimate (12) and after that we apply the panel residual ADF (13) to the estimated residuals of the cointegrating regression (12). To conduct inference with respect to the null hypothesis $\alpha = 0$, we tabulate appropriate critical values through 20,000 trials of the parametric bootstrap procedure described in Appendix B.

Next to a cointegration analysis on the full blown panel, we also test for cointegration in three sub-panels:

- the G7 US dollar exchange rates consisting of Canada, France, Germany, Italy, Japan and the UK (indicated with G7),
- the G10 US dollar exchange rates consisting of the G7 US dollar exchange rates and those of the Netherlands, Sweden and Switzerland (indicated with G10),
- the US dollar exchange rate of Austria plus those of the European Monetary System (EMS) countries France, Germany, Italy, the Netherlands and Spain (indicated with EMS).

The sub-panels were selected on two characteristics: the relative size of the economies and the monetary policy regime. Based on these characteristics we have three sub-panels: one for the G7 countries, one for the G10 countries and one for the original EMS members plus Austria. The first two sub-panels contain countries that yields the bulk of economic activity in our sample of fourteen countries, the last sub-panel contain those countries that have linked their monetary policy to that of Germany either informally (Austria) or formally through the EMS.

In table 7, estimation of (12) for our fourteen dollar-exchange rates yields a cointegrating vector $(\beta_1, \beta_2) = (0.66, -1.34)$. The estimated cointegrating vector for the full panel is quite close to the theoretical valid parameter values although the money elasticity is still rather low compared with the theory. The critical values indicate that this model yields cointegration in the full panel of fourteen dollar-exchange rates at a significance level of 5%. For the sub-panels the results are less positive for our monetary model, as there is only weak evidence for cointegration in the G10-panel. Although the rejection of the null hypothesis of no cointegration is weak for the G10-panel, the corresponding cointegration vector $(\beta_1, \beta_2) = (0.83, -1.52)$ seems to be more in line with the theoretical model than for estimates based on the full panel.

Table 7: Panel cointegration tests for US dollar rates, 1973:1-1994:4[a]

	α	t_α	β_1	β_2
All 14 countries	-0.075	-7.187^{**}	0.664	-1.335
G10 (9)	-0.073	-5.474^{***}	0.832	-1.520
G7 (6)	-0.076	-4.677	0.796	-2.176
EMS (6)	-0.074	-4.281	0.727	-1.813
Critical values t_α:	1%	5%	10%	
All 14 countries	-7.502	-6.917	-6.593	
G10 (9)	-6.395	-5.768	-5.451	
G7 (6)	-5.597	-5.016	-4.715	
EMS (6)	-5.616	-5.026	-4.720	
Power ratios:	1%	5%	10%	
All 14 countries	0.654	0.869	0.931	
G10 (9)	0.445	0.747	0.858	
G7 (6)	0.286	0.580	0.732	
EMS (6)	0.259	0.545	0.695	
1 country	0.022	0.101	0.189	

[a] For α, t_α, β_1 and β_2 see (12) and (13). The number of countries are in parentheses. The critical values of t_α are based on 20,000 trials of our parametric bootstrap procedure (see Appendix B). The symbol $^{*(**)[***]}$ indicates rejection of $\alpha = 0$ at a 1% (5%) [10%] significance level. The power calculations are based on our bootstrap critical values and on the DGP: μ_{it}=$0.9\mu_{it-1}+v_{it}$ (see Appendix B).

As in the case of our individual cointegration tests, we also look at Deutsche Mark numeraire exchange rates in our pooled time series analysis. In analyzing our DM exchange rates, we use the same fourteen exchange rates and corresponding variables as in the case of the individual time series tests. Again we estimate with OLS the pooled model (12) and conduct stationarity tests on the corresponding estimated residuals $\hat{\mu}_{it}$ through FGLS estimation of test equation (13). Like the US dollar panels we set $p = 4$ for all DM exchange rates i. We also analyze the same three sub-panels as in the case of dollar-exchange rates, except that we now look at the exchange rate relationships with respect to Germany.

Using the aforementioned panel cointegration analysis we see in table 8 that the estimated cointegrating vector equals $(\beta_1, \beta_2) = (0.92, -1.82)$ for DM rates. This cointegration vector is more in line with both the theoretical model and the cross-section estimates

Table 8: Panel cointegration tests for DM rates, 1973:1-1994:4[a]

	α	t_α	β_1	β_2
All 14 countries	-0.090	-7.682^{*}	0.923	-1.820
G10 (9)	-0.090	-5.956^{**}	0.904	-2.114
G7 (6)	-0.076	-4.382	0.942	-2.050
EMS (5)	-0.113	-4.931^{**}	1.024	-2.354
Critical values t_α:	1%	5%	10%	
All 14 countries	-7.378	-6.801	-6.493	
G10 (9)	-6.263	-5.643	-5.328	
G7 (6)	-5.597	-4.993	-4.692	
EMS (5)	-5.206	-4.614	-4.305	
Power ratios:	1%	5%	10%	
All 14 countries	0.676	0.875	0.937	
G10 (9)	0.428	0.720	0.836	
G7 (6)	0.232	0.519	0.668	
EMS (5)	0.228	0.512	0.670	
1 Country	0.027	0.110	0.208	

[a] See notes table 7.

in subsection 3.2, than the panel estimates based on dollar-exchange rates. Comparing the empirical t-value for H_0: $\alpha = 0$ with the tabulated critical values, gives a rejection of the null hypothesis at a 1% significance level. For the three sub-panels we can only in the case of the G7-panel not find cointegration based on a theoretical model like (7). Both for the G10-panel as well the EMS-panel we find cointegration at a 5% significance level. In the case of the EMS-panel we see that the estimated cointegrating vector $(\beta_1, \beta_2) = (1.02, -2.35)$ is comparable to the cross-section regression estimates in subsection 3.2 and comply very well with the theory. This result indicates that in a panel of exchange rates which are fixed to the same anchor currency, the adjustment speed to the long-run equilibrium in (7) is quite high. One therefore does not have the need of a large number of cross-sections to find mean reversion based on our monetary model.

We also determine the power of our panel cointegration test with FGLS t_α's to reject the null hypothesis of no cointegration if the alternative of cointegration is true. Power studies are conducted for our four panels and for a hypothetical one-country case. The one-country case is studied to asses whether the failure of cointegration tests for individual countries (section 3.1) is due to a lack of power. We assume that the residuals of (12) are near unit root processes: $\mu_{it} = 0.9\mu_{i,t-1} + \upsilon_{it}$. An AR(1) parameter of 0.9 implies a process which is difficult to distinguish at first sight from a non-stationary process, due to its large persistence. The power calculations are based on the parametric bootstrap procedure which is used to tabulated our critical values, but now with the above mentioned stationary AR(1) process for μ_{it}. A more detailed description can be found in Appendix B.

The power calculations in tables 7 and 8 indicate that in case of a true (though persistent) alternative hypothesis, small panels of five to six countries already give an improved power to reject the null compared to the one-country case. This latter result seems to confirm our assumption that the negative results for individual countries in subsection 3.1 were caused by a lack of power due to a short time span. For all our multi-country samples we see that for 10% and 5% significance levels the power ratios exceed the 50% rate. Only in the case of 1% significance levels we find that we need the full-blown panel of fourteen countries to have powerful inference. Our power analysis also indicate that a failure in certain sub-panels to reject the null of no cointegration can only partly be contributed to moderate power of our tests. Comparing the test results for US dollar and DM panels it is more probable that a failure to reject the null hypothesis for sub-panels is caused by the degree of differences in monetary policy regime within those sub-panels.

5 Conclusions

In this paper we have analyzed a simple, forward-looking monetary exchange rate model based on relative money supplies and relative real incomes. The main objective of our empirical analysis was to establish the appropriateness of our monetary model for long-run dynamics in nominal exchange rates. Therefore the concept of cointegration is central to our approach, as the elements in our model are non-stationary.

As a first step we applied the by now conventional Johansen (1991) cointegration test on each of the fourteen dollar exchange rates in our sample. The result of these individual tests indicate that there is no cointegration based on our monetary exchange rate model. Hence, the long-run explanatory power of the monetary model is very poor for individual exchange rates. In contrast to our individual cointegration tests, regressions based on cross-sections of country averages indicate that our monetary exchange rate model indeed has explanatory power. Not only are cross-section estimates in agreement with the theoretical model, scatter diagrams of partial correlations indicate that the model parameters are fairly homogeneous across our sample of countries.

To enhance the power of our cointegration tests we pool our individual time series into panel data sets. We first estimate our monetary model on our panel data set of fourteen exchange rates, where we restrict the parameters of the relative money supplies and relative real incomes to be identical across our fourteen countries. The parameter estimates in our pooled model are quite close to the theoretical appropriate values. As a second step we test if the residuals of our estimated pooled model are collectively non-stationary, i.e. we test on a pooled time series basis the null hypothesis of no cointegration in our monetary model. The result for our fourteen dollar exchange rates gives a rejection of the null of cointegration on a panel data basis. Results for sub-panels of US dollar exchange rates for G7 and EMS countries are not as positive as for our full-blown panel. Only for a panel of G10 countries can we find weak evidence for cointegration.

To investigate the influence of the choice of numeraire, we applied the previous analysis to the same countries but now for DM exchange rates and fundamentals with respect to Germany. As in the case of dollar exchange rates, cointegration tests for individual countries could not find evidence in favor of our monetary model. If we conduct panel estimates on our sample of DM exchange rates we find that the parameter estimates are very close to the theoretical valid values, more than in the case of dollar exchange rates. We also have a powerful rejection of the null hypothesis of no cointegration. Our cointegration test for sub-panels of DM rates are much more positive for our monetary model than in the case of US dollar rates. For both G10 and EMS sub-panels we find strong evidence for cointegration. Especially in the case of the EMS sub-panel are the results positive: the cointegration vector estimates are fully in-line with the theory and our cross-section estimates.

Power studies for our panel cointegration tests indicate that the negative results for individual countries indeed can be related to a lack of power as a result of a short time span. The power calculations also indicate that rejections of cointegration for certain sub-panels cannot completely be ascribed to a lack of power. The better performance in sub-panels of our model based on DM rates could more probably be related to the greater homogeneity of the monetary policy regimes within the G10 and EMS sub-panels of DM rates. Especially in the DM-panel of EMS countries could this be the case, as all these countries have linked their monetary policies to that of Germany. Therefore one can only observe one monetary policy regime in a DM-panel of EMS countries, namely that of Germany. For the US dollar-panels the diversity of the individual monetary policy regimes is much greater than for DM rates. One has for the US dollar-panels at least two regimes: that of the US and that of Germany.

All in all, our analysis shows that a rational expectations monetary exchange rate model properly describes long-run dynamics in nominal exchange rates. There maybe room for improvement for our theoretical model as different monetary policy regimes have a different impact on money velocity. Disturbances in relative money velocities are in our model the device through which there are deviations with respect to the long-run equilibrium. A better modelling of money velocity could enhance the performance of our model, especially for US dollar exchange rates.

Appendix

A Data sources

We used the *International Financial Statistics* (IFS) of the IMF as our main source of data. In some cases we were forced to use other data sources, as the IFS source could not supply certain data. All the data are on a quarterly basis and start in the first quarter of

1973 (with the breakdown of the Bretton Woods system) and end in the fourth quarter of 1994. The countries that make up our data set are Australia, Austria, Canada, Finland, France, Germany, Italy, Japan, the Netherlands, Norway, Spain, Sweden, Switzerland, United Kingdom and the United States.

Our dollar exchange rates were extracted from line "ae" in the IFS data tape and we constructed our DM exchange rates from these dollar exchange rates through the triangular arbitrage condition. For both numeraire cases we have 14 exchange rates. As our measure of money supply we used the seasonally unadjusted M1 data from the IFS (line 34). Exceptions were Finland, Norway, Sweden and the UK. Finnish M1 is partly obtained from the OECD's *Main Economic Indicators* (MEI) and partly from the IFS for the period before 1980. There is a structural break in 1990 in the Finnish M1 data from the IFS, which is not the case for the MEI data. As the MEI Finnish M1 data are only available from 1980 onwards, we had to link the IFS and MEI series together by multiplying the pre-1980 IFS data with the average of the ratio of MEI and IFS data during the period 1980-1988. Norwegian M1 we completely extracted from the MEI. For Sweden we could not find M1 data and therefore used M2 data from the IFS (as well in that case for the US and Germany to construct the relative money variables). In the case of the United Kingdom we could not find M1 data either and as an alternative we used UK M1 data from De Nederlandsche Bank (the Dutch Central Bank).

As a measure for quarterly real income we used real Gross Domestic Product from the IFS data base (line 99b.c). Again there were exceptions. For the Netherlands there were no GDP data available for the whole sample and as an alternative we used Dutch GDP data from De Nederlandsche Bank. We had the same problem for Germany and Japan, for which we now used real Gross National Product from the MEI data base. All the real income data were seasonally adjusted, with the exception of Austria, Finland, the Netherlands, Norway and Sweden for which we used the X-11 method to make the corresponding series seasonally adjusted in order to get a consistent data set.

B Parametric bootstrap procedures

Like the conventional individual time series ADF on regression residuals, the t-test on $\alpha = 0$ in our panel residual ADF (13) will not have a standard asymptotic t-distribution. To get appropriate critical values for our cointegration test we use a parametric bootstrap to tabulate finite sample critical values under the null hypothesis. Our parametric bootstrap has the following set-up:

1. For each cross-section i we generate artificial fundamental series and residuals μ_{it} as martingales with 188 time series observations. We delete the first 100 observations to correct for any initial-value bias. The remaining 88 time series observations are in compliance with the number of observations for our quarterly sample 1973:1-1994:4. The first differences of each generated martingale comply with the following

processes:

$$\Delta\left(m_{it}-m_{it}^{*}\right)^{s}=\tau_{1,i}\Delta\left(m_{i,t-4}-m_{i,t-4}^{*}\right)^{s}+\pi_{it}^{m}, \tag{14}$$

$$\Delta\left(y_{it}-y_{it}^{*}\right)^{s}=\tau_{2,i}\Delta\left(y_{i,t-1}-y_{i,t-1}^{*}\right)^{s}+\pi_{it}^{y}, \tag{15}$$

$$\Delta\mu_{it}^{s}=\sum_{j=1}^{4}\tau_{3j,i}\Delta\mu_{i,t-j}^{s}+\pi_{it}^{\mu}, \tag{16}$$

where $\tau_{1,i}$, $\tau_{2,i}$, $\tau_{31,i}$, $\tau_{32,i}$, $\tau_{33,i}$ and $\tau_{34,i}$ are the parameters of (14), (15) and (16) estimated on the empirical data for each i separately and the π's are zero mean innovations. A superscript s indicates a generated martingale which act as the artificial counterpart of respectively relative money, relative real income and the residuals of (12) (which are non-stationary under the null). The above mentioned models are in case of $\Delta\left(y_{it}-y_{it}^{*}\right)$ selected through Schwarz's Information Criterion, in case of $\Delta\left(m_{it}-m_{it}^{*}\right)$ to mimic the seasonality in this variable and for $\Delta\mu_{it}$ because we assumed a fourth order serial correlation in μ_{it}. The generated innovations π_{it}^{m}, π_{it}^{y} and π_{it}^{μ} are drawn from a zero mean normal distribution per i, calibrated such that their second moments equal the historical second moments of the residuals of the fitted equivalents of (14), (15) and (16).

2. We allow the artificially generated innovations π_{it}^{μ} for $\Delta\mu_{it}$ to be contemporaneously correlated across i, based on the historical cross-sectional covariance matrix of $\pi_{1t}^{\mu},\ldots,\pi_{Nt}^{\mu}$ in estimates of (16). Through this we mimic the aforementioned cross-section correlation of exchange rate returns.

3. The simulated series for the fundamental variables and the cointegrating regression residuals μ_{it} are used to generate artificial exchange rate data through the DGP:

$$e_{it}^{s}=\hat{\beta}_{0,i}+\hat{\beta}_{1}\left(m_{it}-m_{it}^{*}\right)^{s}+\hat{\beta}_{2}\left(y_{it}-y_{it}^{*}\right)^{s}+\mu_{it}^{s}. \tag{17}$$

   In this DGP $\hat{\beta}_{0,i}$, $\hat{\beta}_{1}$ and $\hat{\beta}_{2}$ are taken from the original estimate of (12) on the empirical data.

4. We pool the generated series of e_{it}^{s}, $\left(m_{it}-m_{it}^{*}\right)^{s}$ and $\left(y_{it}-y_{it}^{*}\right)^{s}$ across the i's and fit regression (12) for these generated series with OLS.

5. We estimate (13) with FGLS for the residuals of (12) estimated on our artificial data (as under item 4) and construct the t-value for H_0: $\alpha=0$.

6. This process (items 1 through 5) is replicated 20,000 times and we use the 1%, 5% and 10% quantiles of the bootstrap distribution of the t-values to tabulate appropriate critical values.

The power analysis of our panel cointegration test is also conducted through a parametric bootstrap procedure. In general this parametric bootstrap is identical to the one which is used to tabulate the critical values for our panel cointegration tests. The difference with the original parametric bootstrap is that we now assume for our power studies that the simulated residuals of (12) are persistent and stationary: $\mu_{it}^{s} = 0.9\mu_{i,t-1}^{s} + \upsilon_{it}$. Next to that we replace the fourth order serial correlation process in (16) with the following process:

$$\upsilon_{it} = \sum_{j=1}^{4} \tau_{3j,i}\upsilon_{i,t-j} + \pi_{it}^{\upsilon}, \tag{18}$$

where $\upsilon_{it} = \mu_{it}^{s} - 0.9\mu_{it}^{s}$ and the $\tau_{3j,i}$'s are the values resulting from estimation of (18) on their empirical equivalents. The historical cross-section covariance matrix of the estimated π_{it}^{υ}'s is used to create cross-section correlation in υ_{it} and through that in Δe_{it}^{s}. In our parametric bootstrap power analysis we use the original critical values, tabulated under the null, to calculate the relative number of rejections of the null hypothesis under the assumption that the alternative of a persistent and stationary AR(1) model in μ_{it} is true. All our power calculations are based on 20,000 simulated panels.

As a benchmark we calculate the power of our residuals-based cointegration test for one country. For the one-country case we also apply parametric bootstraps to calculate both the critical values under the null and the power ratios based on those critical values. The one-country parametric bootstraps are based on the panel estimates of β_1 and β_2 on the panels of fourteen exchange rates. Although these estimates may not be the true cointegration vector for the one-country case, we make use of the result in Engle and Granger (1987) that linear transformations of the original cointegration vector do not influence the statistical properties of cointegration tests. Next to that we use the cross-section means of both the historical second moments of Δe_{it}, $\Delta\left(m_{it} - m_{it}^{*}\right)$, $\Delta\left(y_{it} - y_{it}^{*}\right)$, $\Delta\mu_{it}$, υ_{it} and the empirical estimates of $\beta_{0,i}$, $\tau_{1,i}$, $\tau_{2,i}$ and $\tau_{3j,i}$, for all fourteen exchange rates.

C Cross-section correlation matrices

Table C.1: Cross correlations Δe_{it}, US dollar rates 1973:1-1994:4

	Austr.	Aut.	Can.	Fin.	Fr.	Ger.	Ital.	Jap.	Nl.	Nor.	Sp.	Swed.	Switz.
Austr.	1.00												
Aut.	0.22	1.00											
Can.	0.30	0.05	1.00										
Fin.	0.26	0.76	0.15	1.00									
Fr.	0.25	0.92	0.03	0.73	1.00								
Ger.	0.21	0.99	0.03	0.76	0.92	1.00							
Ital.	0.21	0.79	0.03	0.75	0.84	0.77	1.00						
Jap.	0.35	0.63	0.04	0.45	0.63	0.63	0.58	1.00					
Nl.	0.21	0.99	0.03	0.76	0.92	0.99	0.81	0.64	1.00				
Nor.	0.27	0.89	0.05	0.80	0.85	0.88	0.71	0.56	0.88	1.00			
Sp.	0.24	0.72	0.05	0.71	0.75	0.70	0.77	0.45	0.74	0.73	1.00		
Swed.	0.18	0.77	0.06	0.85	0.75	0.77	0.70	0.45	0.78	0.86	0.74	1.00	
Switz.	0.23	0.88	0.04	0.69	0.84	0.88	0.73	0.68	0.86	0.78	0.59	0.69	1.00
U.K.	0.25	0.68	0.13	0.76	0.71	0.69	0.69	0.56	0.70	0.76	0.67	0.73	0.65

Table C.2: Cross correlations of OLS residuals of (13), US dollar rates 1973:1-1994:4

	Austr.	Aut.	Can.	Fin.	Fr.	Ger.	Ital.	Jap.	Nl.	Nor.	Sp.	Swed.	Switz.
Austr.	1.00												
Aut.	0.17	1.00											
Can.	0.21	0.04	1.00										
Fin.	0.19	0.58	0.06	1.00									
Fr.	0.24	0.86	−0.05	0.61	1.00								
Ger.	0.23	0.91	−0.04	0.62	0.86	1.00							
Ital.	0.19	0.71	0.06	0.61	0.80	0.70	1.00						
Jap.	0.31	0.65	0.18	0.39	0.51	0.57	0.43	1.00					
Nl.	0.11	0.91	0.03	0.54	0.83	0.89	0.70	0.56	1.00				
Nor.	0.32	0.68	0.20	0.59	0.60	0.69	0.54	0.46	0.71	1.00			
Sp.	0.19	0.63	−0.03	0.62	0.66	0.60	0.64	0.44	0.63	0.62	1.00		
Swed.	0.16	0.59	−0.10	0.63	0.57	0.55	0.57	0.30	0.57	0.51	0.56	1.00	
Switz.	0.17	0.81	0.00	0.51	0.80	0.79	0.59	0.62	0.78	0.56	0.49	0.53	1.00
U.K.	0.30	0.56	0.16	0.61	0.61	0.56	0.67	0.47	0.56	0.59	0.58	0.52	0.54

Table C.3: Cross correlations Δe_{it}, DM rates 1973:1-1994:4

	Austr.	Aut.	Can.	Fin.	Fr.	Ital.	Jap.	Nl.	Nor.	Sp.	Swed.	Switz.	UK
Austr.	1.00												
Aut.	0.26	1.00											
Can.	0.79	0.25	1.00										
Fin.	0.56	0.25	0.63	1.00									
Fr.	0.34	0.31	0.31	0.31	1.00								
Ital.	0.38	0.38	0.42	0.55	0.59	1.00							
Jap.	0.57	0.16	0.51	0.29	0.30	0.38	1.00						
Nl.	0.18	0.48	0.20	0.26	0.31	0.50	0.22	1.00					
Nor.	0.55	0.31	0.56	0.63	0.35	0.31	0.31	0.18	1.00				
Sp.	0.47	0.36	0.49	0.57	0.48	0.61	0.29	0.44	0.53	1.00			
Swed.	0.35	0.21	0.42	0.72	0.28	0.40	0.18	0.23	0.68	0.57	1.00		
Switz.	0.11	0.10	0.07	0.10	0.18	0.18	0.31	−0.04	0.05	−0.02	0.07	1.00	
UK	0.49	0.12	0.54	0.67	0.39	0.50	0.44	0.30	0.62	0.54	0.56	0.15	1.00
US	0.77	0.23	0.95	0.61	0.33	0.44	0.53	0.21	0.58	0.51	0.43	0.06	0.53

Table C.4: Cross correlations of OLS residuals of (13), DM rates 1973:1-1994:4

	Austr.	Aut.	Can.	Fin.	Fr.	Ital.	Jap.	Nl.	Nor.	Sp.	Swed.	Switz.	UK
Austr.	1.00												
Aut.	0.28	1.00											
Can.	0.68	0.46	1.00										
Fin.	0.41	0.24	0.51	1.00									
Fr.	0.32	0.45	0.34	0.37	1.00								
Ital.	0.39	0.41	0.59	0.51	0.61	1.00							
Jap.	0.53	0.54	0.60	0.39	0.36	0.41	1.00						
Nl.	0.26	0.54	0.42	0.13	0.36	0.42	0.35	1.00					
Nor.	0.37	0.31	0.45	0.39	0.12	0.33	0.35	0.31	1.00				
Sp.	0.41	0.41	0.50	0.54	0.41	0.55	0.46	0.35	0.45	1.00			
Swed.	0.42	0.35	0.41	0.41	0.29	0.45	0.25	0.28	0.31	0.50	1.00		
Switz.	0.12	0.37	0.13	0.10	0.42	0.19	0.38	0.31	0.09	0.13	0.24	1.00	
UK	0.47	0.35	0.58	0.51	0.48	0.62	0.49	0.40	0.36	0.52	0.50	0.29	1.00
US	0.69	0.44	0.87	0.52	0.44	0.56	0.56	0.44	0.33	0.56	0.48	0.17	0.59

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