Deutsches Institut für
Wirtschaftsforschung

DIW BERLIN

Discussion Papers

1119

Joscha Beckmann • Ansgar Belke • Frauke Dobnik

Cross-section Dependence and the Monetary Exchange Rate Model A Panel Analysis

Berlin, April 2011

IMPRESSUM

DIW Berlin
German Institute for Economic Research
Mohrenstr. 58
10117 Berlin

Tel. +49 (30) 897 89-0
Fax +49 (30) 897 89-200
http://www.diw.de

ISSN print edition 1433-0210
ISSN electronic edition 1619-4535

Papers can be downloaded free of charge from the DIW Berlin website:
http://www.diw.de/discussionpapers

Discussion Papers of DIW Berlin are indexed in RePEc and SSRN:
http://ideas.repec.org/s/diw/diwwpp.html
http://www.ssrn.com/link/DIW-Berlin-German-Inst-Econ-Res.html

Cross-section Dependence and the Monetary Exchange Rate Model – A Panel Analysis[1]

Joscha Beckmann[a], Ansgar Belke[a,b,c,*], Frauke Dobnik[a]

[a] *Department of Economics, University of Duisburg-Essen, Universitätsstr. 12, 45117 Essen, Germany*

[b] *German Institute for Economic Research (DIW Berlin), Mohrenstrasse 58, 10117 Berlin, Germany*

[c] *Institute for the Study of Labour (IZA), Schaumburg-Lippe-Strasse 5-9, 53113 Bonn, Germany*

April 12, 2011

Abstract. This paper tackles the issue of cross-section dependence for the monetary exchange rate model in the presence of unobserved common factors using panel data from 1973 until 2007 for 19 OECD countries. Applying a principal component analysis we distinguish between common factors and idiosyncratic components and determine whether non-stationarity stems from international or national stochastic trends. We find evidence for a cross-section cointegration relationship between the exchange rates and fundamentals which is driven by those common international trends. In addition, the estimated coefficients of income and money are in line with the suggestions of the monetary model.

JEL-Classification: C32, C23, F31, F41

Keywords: Monetary exchange rate model, common factors, panel data, cointegration, vector error-correction models

[*]Corresponding author. Tel.:+49 (0)201 183-2277; fax:+49 (0)201 183-4181.
E-mail addresses: joscha.beckmann@uni-due.de (J. Beckmann), ansgar.belke@uni-due.de (A. Belke) and frauke.dobnik@uni-due.de (F. Dobnik).
[1]We are grateful for research assistance from Alexander Haering.

1 Introduction

The question whether exchange rates are cointegrated with fundamental factors is still a controversial research area in economics. On a country base, the results crucially depend on the sample and the countries under investigation. Although fundamental factors suggested by the monetary model have mostly been unsuccessful at forecasting exchange rates, a result first highlighted in the seminal study by Meese and Rogoff (1983), many studies have found evidence of a long-run relationship between exchange rates and fundamentals when more sophisticated econometrics such as panel methods are applied.[1]

In the context of market efficiency and exchange rates, another strand of literature focuses on the question whether major exchange rates share common stochastic trends or, more precisely, whether co-movements between exchange rates can be identified. Evidence for this kind of relationships has been found both before and after the introduction of the Euro by different authors (see, e.g., Haug, MacKinnon, and Michelis (2000); Kühl (2010)).[2] Kühl (2008) shows that not only exchange rates share common stochastic trends but also cointegration between fundamentals across the economies exists. For instance, cointegration across countries might occur if monetary policies are coordinated to limit exchange rate fluctuations such that currency prices cannot permanently diverge from each other (Phengpis and Nguyen, 2009).

[1] In the time series dimension these econometrics include non-linear approaches such as non-linear error correction models (see, e.g., Taylor and Peel (2000); Taylor, Peel, and Sarno (2001)) or models with time varying coefficients (see, e.g., Frömmel, MacDonald, and Menkhoff (2005a,b); Goldberg and Frydman (2001) and Yuan (2011)).

[2] Although Granger (1986) raised the argument that cointegration between two or more asset prices violates the weak form of market efficiency due to the predictability of asset prices based on the past prices of other assets, it is controversially discussed whether cointegration between exchange rates actually implies market inefficiency (see, e.g., Phengpis and Nguyen (2009)).

Evident, the issue of co-movements is of crucial importance when more than one economy is analysed simultaneously. Compared to country by country studies, panel data analyses of the monetary model have the advantage to increase the sample size and ceteris paribus lead to more precise estimates. However, most panel data studies assume that cross-section dependencies between countries do not exist, a condition which is very likely to be violated in reality of empirical work. Exactly this deficiency has recently been emphasised by Basher and Westerlund (2009) who base their analysis of the monetary model on the dataset of the influential study by Mark and Sul (2001). Their results suggest that accounting for the effects of cross-section dependence is crucial when analysing the monetary model in the panel context. The reason is that the monetary model is more likely to hold when those effects are considered. However, the authors do not test for cross-section cointegration. Further, they do not estimate the coefficients of the monetary model explicitly. Previous authors who found evidence for the monetary model using panel data also have not paid attention to the issue of cross-section cointegration.

One of the first studies which analyses a panel in the context of empirical tests of the monetary exchange rate model is Husted and MacDonald (1998). They examine three panel datasets constructed for the US dollar, the Deutschmark and the Japanese Yen and provide evidence of cointegration relationships between the exchange rates and fundamentals in all cases. Groen (1999) studies a panel of US dollar nominal exchange rates for 14 industrialised countries between 1973 and 1994. His coefficient estimates are mostly consistent with the monetary model for his full panel and three sub-panels. What is more, Mark and Sul (2001) analyse the long-run relationship between the nominal exchange rate and fundamentals for a panel of quarterly data for 19 OECD countries from 1973 until 1997. By allowing heterogeneous short-run dynamics the authors provide further evidence that the nominal exchange rate is cointegrated with monetary fundamentals. The framework applied by Groen (2002) allows to test for a joint number of cointegration vectors per country where the error-correction estimates are assumed to be heterogeneous. Besides delivering estimates which are in line with the monetary model the results also show that panel-based cointegration techniques are more powerful in case of low mean reversion and short span of data. In a critical evaluation, Rapach and Wohar (2004) compare the performance of the monetary model on a country by country basis with results based on panel analysis for the same dataset as Mark and Sul (2001). They conclude that while pooling the data increases the sample size as well as the support for the monetary model the risk of obtaining spurious evidence of cointegration also rises when panel tests are applied. They identify the assumption of a common data generating process and homogeneity restrictions as important caveats. Recently, Cerra and Saxena (2010) exploit the power of panel cointegration tests by including a broad country sample of 98 countries with annual data spanning from 1960 until 2004. Their results provide further evidence that monetary fundamentals play an important role for the nominal exchange rate.[3]

Summing up the literature, the issue of cross-section dependence has been addressed by allowing for a common deterministic time trend or applying seemingly unrelated regression (SUR) estimates which account for contemporaneous correlation between the errors across equations by some studies. However, the possibility of cross-section cointegration which arises in case of common stochastic trends across countries has been neglected. In general the application of panel tests in

[3]Many other panel studies have focused on the validity of the purchasing power parity (PPP) for more than one country by testing the hypothesis of stationary real exchange rates. Such studies have for example been carried out by Hakkio (1984), Abuaf and Jorion (1990) and Wu (1996). See Sarno and Taylor (2003) for an overview.

the presence of cross-section cointegration can lead to biased conclusions (Banerjee, Marcellino, and Osbat, 2004). Hence, we focus on cross-section dependence in terms of common stochastic trends rather than correlations between errors across panel members since the latter does not necessarily imply cointegration across those members (Breitung and Pesaran, 2008). Accordingly, the question arises whether the long-run relationship between exchange rates and fundamentals which has been identified by many studies mentioned above is mainly driven by such a cointegration relation across countries. To tackle this issue in consideration of the monetary model we pay special attention to the role of strong cross-section dependence. As suggested by Breitung and Pesaran (2008) strong dependence arises when there are unobserved common factors which are able to identify common stochastic trends. For this purpose, we apply a principal component analysis (PCA) as proposed by Bai and Ng (2004) to distinguish between common factors and idiosyncratic components. Additionally, we extend the dataset of Mark and Sul (2001) until the end of 2007 to account for recent developments. By applying PCA the resulting idiosyncratic component can be interpreted as that part of the variable which is driven by national trends while the common component, in contrast, represents international trends in the evolution of the variable. Thus, we are also able to assess whether the non-stationarity of nominal exchange rates and fundamental factors as well as a long-run relationship between both are mainly driven by international trends or national developments. In this respect we follow Dreger (2010) who applies a similar approach to account for cross-section dependence when analysing the real interest rate parity (RID) condition for different sub-periods.

Our analysis fills a gap in the literature in several respects. If cross-section cointegration is found, the long-run relationship can be considered to be driven mainly by common stochastic trends, for example by an international business cycle or global excess liquidity. This would imply that a coordination of monetary or exchange rate policies is more likely to succeed. If, by contrast, cointegration is found to be due to idiosyncratic components, national policies need to account for specific national developments when making their decisions. The distinction between national and international shocks is even more important, for instance, if countries decide to join a currency union. A common currency and a unified monetary policy face severe difficulties if the members are confronted with frequent and/or huge asymmetric shocks.

The remainder of this paper is organised as follows: In section 2 we provide a brief overview of the model we consider in our analysis. Section 3 expounds the problem of cross-section dependence, discusses the econometric methodology to take account for that and describes the data. Section 4 presents the results of our empirical work. Section 5 concludes.

2 Monetary model of the exchange rate

After the breakdown of Bretton Woods, the monetary approach emerged as the most popular exchange rate model. Monetary models of the exchange rate rely on purchasing power parity (PPP) as a long-term equilibrium of good markets. With respect to asset markets, the non-arbitrage condition of the uncovered interest rate parity (UIP) is assumed to hold. As a starting point, consider the following money demand function of the form

$$m_{it} - p_{it} = \gamma_i + \phi_i y_{it} - \lambda_i i_{it} + \mu_{it}, \tag{1}$$

where the index $i = 1, ..., N$ represents the countries in the panel and $t = 1, ..., T$ refers to the time period. A country-specific intercept is denoted by γ_i and m_{it}, p_{it} and y_{it} are the logarithm of money supply, price level and real income, respectively. The interest rates, i_{it}, are expressed in percentage. Thus, ϕ_i and λ_i are measures of income and interest rate elasticity. By assumption, a similar equation holds abroad. Further, it is necessary that the PPP holds for each country

$$p_{it} = \nu_i + p_t^f + s_{it} + e_{it}, \tag{2}$$

where the superscript f denotes to the foreign country and s_{it} is the logarithm of the nominal exchange rate between the domestic and foreign country. Next, taking the difference between the domestic and foreign money demand function and substituting the price differential by the nominal exchange rate according to PPP in equation (2) gives

$$s_{it} = \alpha_i + (m_{it} - m_t^f) - \phi_i(y_{it} - y_t^f) + \lambda_i(i_{it} - i_t^f) + u_{it} \tag{3}$$

where $u_{it} = \mu_t^f - \mu_{it} - e_{it}$ and $\alpha_i = \gamma^f - \gamma_i - \nu_i$ which denotes a constant term that is zero in the original model. A rise of the exchange rate s_{it} corresponds to a depreciation of the domestic currency. If the uncovered interest rate parity (UIP) holds $(i_{it} - i_t^f)$ can be replaced by the expected change in the exchange rate $E_t(s_{t+1}) - s_t$[4]

$$s_{it} = \alpha_i + (m_{it} - m_t^f) - \phi_i(y_{it} - y_t^f) + \lambda_i(E_t(s_{t+1}) - s_t) + u_{it}. \tag{4}$$

[4]With an expectation-generating mechanism based upon PPP the differences in interest rates can then be replaced by the differences in expected rates of inflation. The latter, in turn, is substitutable by the actual inflation differential. The real interest rate model (RID) by Frankel (1979) can be derived by combining the resulting equation with UIP and condition an expectation formulation where the expected rate of depreciation is a function of an equilibrium exchange rate and the expected long-run inflation differential (Frankel, 1979).

The simplest form of the flexible price-monetary approach arises if the expected change in the exchange rate is considered to be stationary. With relative monetary velocity taken as a measure of monetary fundamentals, the nominal exchange rate is then determined by the difference between domestic and foreign money supply and the proportion of income between both economies (Frankel (1979); Taylor and Peel (2000)):

$$s_{it} = \alpha_i + (m_{it} - m_i^f) - \phi_i(y_{it} - y_i^f) + \varepsilon_{it}, \tag{5}$$

where $\varepsilon_{it} = \lambda_i(E_t(s_{i+t}) - s_t) + u_{it}$. An increase of the domestic money supply leads to excess money supply and consequently results in an increase of domestic prices to restore money market equilibrium. As purchasing power parity (PPP) is assumed to hold, the domestic currency depreciates as a result of the rise in prices. In case of a domestic income expansion, money demand increases, domestic prices fall and the currency appreciates. Applying this formulation of the monetary model in our analysis, we follow most panel studies such as Groen (1999, 2002), Rapach and Wohar (2004) and Basher and Westerlund (2009) who base their empirical work on the same framework.

3 Econometric methodology and data

3.1 Cross-section dependence

It is widely known that standard unit root and cointegration tests based on individual time series have low statistical power, especially when the time series is short (Campbell and Perron, 1991). In contrast, panel data methods have greater power by extending the time series dimension by the cross-sectional dimension, allowing for higher degrees of freedom. As panel-based tests rely on a broader information set, the power can substantially be increased and tests are more accurate and reliable. However, first generation panel unit root and cointegration tests have been heavily criticised because they assume that the cross-section members are independent. This condition is often likely to be violated, for example, because common stochastic trends may occur due to global developments or strong relationships between economies. The main reason why residual based tests use the assumption of cross-sectional independence is the fact that standard asymptotic tools, such as the Central Limit Theorem, can be applied in this case. Inappropriately assuming cross-sectional independence when cross-section cointegration is present can, however, distort the panel results (Banerjee et al. (2004), Urbain and Westerlund (2006)). Using simulation methods, Banerjee et al. (2004) show that neglecting cross-section cointegration, such as previous panel studies, has important distortionary effects.

In the present study, the possibility of cointegration across countries when testing for the monetary model is taken into account by applying the no-cointegration test approach suggested by Gengenbach, Palm, and Urbain (2006). Their sequential testing strategy is based on the following common factor structure to model cross-section dependence[5]

$$Y_{i,t} = \theta_{1i} F_{1t} + E_{1i,t}, \quad \text{and} \tag{6}$$

$$X_{i,t} = \theta_{2i} F_{2t} + E_{2i,t}, \tag{7}$$

[5]A different approach is emphasised by Banerjee et al. (2004) who suggest testing for the presence of cross-unit cointegration based on a unit-by-unit cointegration analysis.

where the index $i = 1, ..., N$ represents the cross-sections and $t = 1, ..., T$ refers to the time period. F denotes the common factors and E stands for the idiosyncratic components of the respective variable. A model with N countries leads to N idiosyncratic components for each variable but contains only a small number of common factors. Gengenbach et al. (2006) propose the application of the Bai and Ng (2004) PANIC methodology to test these two uncorrelated components separately for unit roots instead of testing the original variables $Y_{i,t}$ and $X_{i,t}$. Accordingly, this approach allows to determine whether non-stationarity stems from a pervasive or an idiosyncratic source. Further, a cointegration relationship between the original variables $Y_{i,t}$ and $X_{i,t}$ based on the factor structure under equations (6) and (7)

$$Y_{i,t} - \beta_i X_{i,t} = \theta_{1i} \left(F_{1t} - \beta_i \frac{\theta_{2i}}{\theta_{1i}} F_{2t} \right) + E_{1i,t} - \beta_i E_{2i,t} \tag{8}$$

requires that both the common and the idiosyncratic parts of the error term are stationary (see equation (8)). Therefore, Gengenbach et al. (2006) consider two relevant cases. First, the common factors are I(1), while the idiosyncratic components are I(0). In this case non-stationarity in the panel is solely driven by a reduced number of common stochastic trends. Hence, a cointegration relationship between $Y_{i,t}$ and $X_{i,t}$ can occur only if the common factors of $Y_{i,t}$ cointegrate with those of $X_{i,t}$, meaning the existence of cross-section cointegration. The null hypothesis of no-cointegration between the common factors can be investigated using standard time series tests such as the Johansen reduced rank approach (Johansen, 1995). The second case proposed by Gengenbach et al. (2006) refers to the situation in which both common and idiosyncratic stochastic trends are present in the data. In this case, both the common factors and the idiosyncratic components are I(1) and have to be tested separately for cointegration. Since the defactored series are independent by construction, cointegration between the idiosyncratic components can be explored by first generation panel cointegration tests such as those of Pedroni (1999, 2004).

3.2 Data

We extend the Mark and Sul (2001) quarterly dataset for nominal exchange rates relative to the US, nominal money supply, industrial production, and prices which starts in 1973 until the end of 2007. Overall, our sample includes 19 OECD countries: Australia, Austria, Belgium, Canada, Denmark, Finland, France, Germany, Great Britain, Greece, Italy, Japan, Korea, the Netherlands, Norway, Spain, Sweden, Switzerland, and the United States which serve as a numéraire. Similar to the data provided by Mark and Sul (2001), the additional data is taken from International Financial Statistics of the IMF. The nominal money supply is the sum of money and quasi-money for most countries. We use industrial production as a measure for income since real GDP is unavailable for a number of countries. For more details on the data see Mark and Sul (2001).

The further proceeding according to Gengenbach et al. (2006) is as follows. As a first step, we decompose the nominal exchange rate, money supply and income into the two uncorrelated components. As a second step, we test both the common factors and the idiosyncratic components of each variable separately for unit roots and cointegration relationships.

4 Empirical results

4.1 Variable decomposition

Following the Bai and Ng (2004) PANIC methodology as proposed by Gengenbach et al. (2006) the starting point of our empirical investigation of the monetary exchange rate model is the decomposition of each variable into the two uncorrelated components, i.e. a common and an idiosyncratic component. The idiosyncratic component is a residual, which captures the impact of shocks affecting the respective variable of an individual country. These country-specific shocks, for instance, domestic money supply shocks, may have large but geographically concentrated effects. The common component of a variable is 'common' in the sense that it depends on a small number of common shocks, which affect the respective variable of all the countries. We use the principal component analysis for decomposition to obtain consistent estimates of the common factors. Because of potential non-stationarity of the factors we take differenced data, as proposed by Bai and Ng (2004). After estimating the common factors we re-cumulate them to match the integration properties of the original variables. The idiosyncratic components are obtained from a regression of the original series on their common factors. For the money and income variables, we carry out the decomposition before putting it into proportion to the US quantities. This seems reasonable because taking the difference first and decomposing afterwards probably would produce biased results with the common component mainly mirroring movements of the US quantities.

For each variable, two common components are enough to capture at least one third of the overall variance. Any further component would raise the cumulative proportion of the variance only slightly and preliminary evidence shows that results do not qualitatively change.

4.2 Unit root tests

Applying unit root tests to the common factors of the nominal exchange rate and to the common factors of income and money supply relative to the US is important as there is some evidence that some economic variables like money supplies might be better approximated as I(2) rather than I(1) (Juselius, 2007). On the contrary, the common factors might also be stationary if the non-stationarity of the original variable is mainly driven by the idiosyncratic component. To test the null hypothesis of a unit root, we apply standard time series tests, i.e. the augmented Dickey and Fuller (1979) (ADF) test and the Phillips and Perron (1988) (PP) test. As selection rule for the lag order in the ADF regressions we apply the Modified Schwarz Information Criterion (MSIC) proposed by Ng and Perron (2001). The MSIC takes into account that the bias in the estimate of the sum of the autoregressive coefficients is highly dependent on the truncation lag by using a penalty factor that is sample dependent. This modification is evidently more robust when there are negative moving-average errors, which is a fairly common occurrence in macro time series data.[6] According to the results displayed in Table 1, the common factors of all three variables turn out to be non-stationary whereas they become stationary by taking first differences. Hence, the results provide evidence that all common factors are integrated of order one, i.e. I(1). This finding, in turn, allows for the possibility that a cointegration relationship between the common factors of the nominal exchange rates and the common factors of the fundamentals relative to the US exists.

Stochastic trends in the idiosyncratic components can be efficiently explored by first

[6]In contrast, selection rules such as the Schwarz Information Criterion and the Akaike Information Criterion tend to select a lag length that is too small for unit root tests to have a good size.

Table 1: Time series unit root tests for the common components

Variable	Levels				Differences	
	ADF	PP	ADF	PP	ADF	PP
	without trend		*with trend*			
s^c	-1.61(0)	-2.07[6]	-0.99(0)	-1.69[6]	-4.27(2)***	-9.95[4]***
$m^c - m^{US}$	-1.21(2)	-0.69[9]	-1.37(10)	-3.64[7]**	-1.65(4)*	-3.69[0]***
$y^c - y^{US}$	-2.15(0)	-2.15[0]	-2.14(0)	-2.14[0]	-6.11(1)***	-7.27[8]***

Notes: The superscript c denotes the common factor of the respective variable. Numbers in parentheses are the maximum numbers of lag determined by empirical realisations of the Modified Schwarz Information Criterion. Numbers in brackets represents the automatic Newey-West bandwidth selection using the Bartlett kernel. ***, ** and * indicate significance at the 1%, 5% and 10% levels.

generation panel unit root tests, since the defactored series are independent by construction and, thus, fulfil the assumption of cross-sectional independence. In this study we apply the Levin, Lin, and Chu (2002) (LLC) test, the Fisher-type ADF test and the Fisher-type PP test (see Maddala and Wu (1999) and Choi (2001)). The LLC test assumes a homogenous autoregressive parameter for all cross-sections under the alternative whereas the non-parametric Fisher-type tests allow that there are some cross-section units without a unit root.[7] In contrast to the time series unit root evidence for the common components, the results of the panel unit root tests suggest that the idiosyncratic components of the variables under investigation are widely stationary (see Table 2).

Hence, the results indicate that the non-stationarity in the nominal exchange rate, money supply and income relative to the US of the 18 economies are driven mainly

[7]Conducting further panel unit root tests would certainly make sense. However, considering that the idiosyncratic components are residuals by definition we neither include a trend nor a constant such that the analysis is restricted to those tests mentioned above.

Table 2: Panel unit root tests for the idiosyncratic components

Variable	LLC	ADF-Fisher	PP-Fisher
s^i	-2.95***	59.84***	65.97***
m^i	-3.52***	44.24	47.82*
y^i	-4.56***	75.51***	82.27***

Notes: The superscript *i* denotes the idiosyncratic component of the respective variable. Probabilities for the Fisher tests are computed using an asymptotic Chi-square distribution. The LLC test assumes asymptotic normality. The choice of lag levels for the Fisher-ADF test is determined by empirical realisations of the Modified Schwarz Information Criterion. The LLC and Fisher-PP tests were computed using the Bartlett kernel with automatic bandwidth selection. ** and * indicate significance at the 1% and 5% levels.

by common stochastic trends rather than country-specific developments.[8] As a consequence, a long-run equilibrium relationship between exchange rates and fundamentals may exist between the common rather than the idiosyncratic components, which would be equivalent with cross-section cointegration.

4.3 Cointegration analysis

As integration of order one for the common factors and stationarity for the idiosyncratic components are established, the next step is to determine whether both cointegration between the common factors, i.e. cross-section cointegration, and, consequently, cointegration between the underlying variables (see equation (8)) can be verified. The existence of a long-run relationship between the common factors can be investigated using standard time series tests such as the Johansen reduced rank approach (Johansen, 1995).[9] As mentioned before, a small sample size can induce

[8] The common factors of money supply and income are also non-stationary if the unit roots tests are applied without subtracting the US quantities.

[9] The idea of the test is to separate the eigenvalues $\lambda_i, i = 1, ..., r$ which correspond to stationary relations from those eigenvalues $\lambda_i, i = r + 1, ..., p$ which belong to non-stationary eigenvectors. The test statistic of the corresponding likelihood test, the so called trace test, is given by $trace(r) =$

biased realisations of the Johansen test statistics. Hence, we follow Reinsel and Ahn (1992) and Reimers (1992) by applying a modification of the test statistics to account for potential small sample bias. Accordingly, we multiply the Johansen statistics with the scale factor $(T - pk)/T$, where T is the number of observations, p the number of variables and k the lag order of the VAR, such that a proper inference can be made even if the sample size is small. The empirical realisations of both the modified Johansen trace statistic and those of the modified Johansen maximum eigenvalue statistic provide evidence in favour of a long-run relationship between the common factors of nominal exchange rate, money supply and income (see Table 3). Considering that our analysis is based on common factors, this result suggests that cross-section cointegration is existent and important to incorporate when analysing the monetary model. What is more, according to equation (8) both the common and the idiosyncratic part are evidently stationary as required for a cointegration relationship of the underlying variables. Hence, we find support for the monetary exchange rate model.

Table 3: Results of Johansen's tests for cointegration among common components

H_0	Trace Statistic	Critical Value	Max. Eigenvalue Statistic	Critical Value
None	54.96*	42.92	34.16*	25.82
At most 1	20.80	25.87	15.80	19.39
At most 2	5.00	12.52	5.00	12.52

Notes: Potential small sample bias is corrected by multiplying the Johansen statistics with the scale factor $(T - pk)/T$, where T is the number of observations, p the number of variables and k the lag order of the underlying VAR model in levels, see Reinsel and Ahn (1992) and Reimers (1992). Critical values are taken from MacKinnon et al. (1999), and are also valid in case of the small sample correction. The choice of a lag level of two is determined by the Schwarz Information Criterion. A * indicates the rejection of the null hypothesis of no cointegration at least at the 5% level.

As a next step, we explicitly estimate the long-run coefficients of the established

$-(T - p) \sum_{i=r+1}^{N} log(1 - \lambda_i)$.

cointegration relationship between the common factors relative to the US using the dynamic ordinary least squares (DOLS) estimator proposed by Mark and Sul (2003). The DOLS estimator conveniently corrects standard OLS for any bias which might be induced by endogeneity and serial correlation. First, we regress the endogenous variable in each equation on the leads and lags of the first-differenced regressors from all equations to control for potential endogeneities. Next, we apply the OLS method using the residuals from the first step regression. Harris and Sollis (2003) suggest that parametric approaches such as DOLS are more robust than non-parametric if the data have significant outliers and also have less problems in cases where the residuals have large negative moving average components, which is a fairly common occurrence in macro time series data. Hence, we use DOLS to estimate the following model:

$$s_t^c = \alpha + \delta t + \beta_1(m_t^c - m_t^{US}) + \beta_2(y_t^c - y_t^{US}) + \varepsilon_t, \tag{9}$$

where the subscript $t = 1, ..., T$ refers to the time period and t represents a deterministic time trend. The superscript c denotes the common factors of the original variables. Since all variables are specified in natural logarithms, the estimated long-run coefficients can be interpreted as elasticities.

The elasticities of income and money supply turn out to be highly significant and show signs which are in line with the suggestions of the monetary model described in section 2 ($\beta_1 = 0.20$ [$t = 4.09$], $\beta_2 = -1.60$ [$t = -5.18$]). Considering that our results refer to the common factors net of the idiosyncratic component of the underlying variables, our results indicate that an increase in money supply relative to the US results in a depreciation against the dollar while the opposite is true if income relative to the US increases. The fact that the estimated income elasticity is higher than the estimated elasticity of money supply is in line with the results of previous studies by Groen (1999, 2002), Mark and Sul (2001) and Rapach and Wohar (2004). A possible explanation could be the Balassa-Samuelson effect which indicates that a larger income elasticity is a result of the influence of production differentials between countries on real exchange rates (Groen (2002)). Given the fact that the established cross-section cointegration indicates that the countries under investigation share a common stochastic trend, the global business cycle as well as the overall direction of monetary policies seem to be of long-run relevance for movements of the dollar exchange rates. Applying an ADF unit root cointegration test based on the corrected P-value provided by MacKinnon (1996) verifies the stationarity of the established cross-section cointegration relationship between the exchange rate and the fundamental factors ($t = -2.12$ [0.03]).

Additionally, we change our sample in order to test for the reliability of our overall results. We conduct estimations which start in 1976 after the end of the major turbulences from the first oil price shock and also end our analysis prior to the introduction of the euro for both starting dates. In all cases, the overall empirical results of the unit root tests, the cointegration test and the long-run coefficients remain unchanged. The results are available upon request from the authors.

4.4 A dynamic panel error-correction model

In addition to the estimation of the long-run relationship between the common factors of the exchange rates and the fundamentals, we further analyse whether the domestic exchange rates readjust towards this established common international equilibrium relation after a shock occurs. For this purpose, we estimate a dynamic panel-based error-correction model using a two-step procedure and allowing for heterogeneous error-correction behaviour. First, we employ the long-run equation specified in (9) to obtain the deviation from the established long-run equilibrium of the common factors, i.e. $\varepsilon_{i,t}$. Then we estimate the error-correction model with the original variables including both the common and the idiosyncratic components and incorporate the one-period lagged residual from the first step as dynamic error-correction term:

$$\Delta s_{i,t} = \alpha_i + \sum_{k=0}^{p-1} \gamma_{1i,k} \Delta(m_{i,t-k} - m_{t-k}^{US}) + \sum_{k=0}^{q-1} \gamma_{2i,k} \Delta(y_{i,t-k} - y_{t-k}^{US}) + \sum_{k=1}^{r-1} \gamma_{3i,k} \Delta s_{i,t-k} + \lambda_i \varepsilon_{i,t-1} + u_{i,t}, \tag{10}$$

where Δ denotes the first-difference operator, λ_i represents the adjustment coefficient and $u_{i,t}$ is the serially uncorrelated error term with mean zero. We select the lag lengths p, q and r by applying the general-to-specific methodology and, hence, we start with a fairly general specification and exclude most insignificant variables step by step. Since the sample under investigation includes more than 130 observations, the usual finite sample bias of dynamic panel estimations, the so-called Nickell-bias (Nickell, 1981), should be negligible. Consequently, the use of an instrument estimator such as the GMM estimator proposed by Arellano and Bond (1991) is not required. Instead, we use the seemingly unrelated regression (SUR) method to account for heteroscedasticity and contemporaneous correlation in the errors across equations, $u_{i,t}$ and $u_{j,t}$, $i \neq j$. Applying the SUR method, we estimate the parameters of the system (10) by feasible generalised least squares (FGLS). In our context, it is of particular interest whether the domestic nominal exchange rates converge to the established common equilibrium path. These long-run dynamics can be studied by testing the significance of the adjustment coefficients λ_i. Table 4 shows the estimated coefficients and the corresponding P-values of the panel-based error-correction model.

According to our estimations, the adjustment coefficients of all countries, except Greece, turn out to be small but highly significant and correctly signed. Our result suggests that the national exchange rates indeed adjust to disequilibria from the established long-run relation. This finding accessorily highlights the relevance of the common cross-section cointegration relationship from a domestic point of view.

5 Conclusions

Applying the no-cointegration test approach suggested by Gengenbach et al. (2006) including a principal component analysis and extending the Mark and Sul (2001) quarterly dataset, we have shed some light on the source of non-stationarity in the exchange rates and fundamentals. While previous panel studies on the monetary exchange rate model have been subject to critique because they arguably neglect cross-section dependence in general or cross-section cointegration in particular, our framework is able to deal with this issue by assuming a common factor structure. In this vein, our results suggest that common international rather than national stochastic trends are responsible for the non-stationarity of exchange rates and fundamentals. Further, we are able to show that cross-section cointegration exists in the sense of a long-run relationship between the common factors of exchange rates and fundamentals. The pattern of these results is in line with previous studies by Haug et al. (2000), Phengpis and Nguyen (2009) and Kühl (2010) which report cointegration between exchange rates across countries.

Using data coded relative to the US economy, our findings imply that the dollar

Table 4: Short-run dynamics and adjustment coefficients of the exchange rates of 18 OECD countries

Independent	SUR system estimated by FGLS			
Variables	Coefficient	P-Value	Coefficient	P-Value
	Country-specific intercept		Country-specific adjustment coefficient	
Australia	0.042	0.00	-0.048	0.01
Austria	0.097	0.00	-0.037	0.00
Belgium	0.126	0.00	-0.035	0.00
Canada	0.042	0.00	-0.056	0.00
Denmark	0.091	0.00	-0.039	0.00
Finland	0.101	0.00	-0.047	0.00
France	0.063	0.00	-0.037	0.00
Germany	0.024	0.00	-0.035	0.00
Greece	0.015	0.55	0.001	0.89
Italy	0.344	0.00	-0.056	0.00
Japan	0.132	0.00	-0.034	0.00
Netherlands	0.037	0.00	-0.035	0.00
Norway	0.082	0.00	-0.035	0.00
South Korea	0.254	0.00	-0.040	0.00
Spain	0.168	0.00	-0.038	0.00
Sweden	0.081	0.00	-0.035	0.00
Switzerland	0.034	0.00	-0.043	0.00
United Kingdom	0.017	0.00	-0.029	0.01
	Short-run dynamics			
$\Delta m_{i,t}$	-0.017	0.36		
$\Delta y_{i,t}$	-0.034	0.01		
$\Delta m_{i,t-1}$	-0.042	0.02		
$\Delta y_{i,t-1}$	0.035	0.02		
$\Delta s_{i,t-1}$	0.035	0.09		
$\Delta y_{i,t-2}$	0.056	0.00		
$\Delta s_{i,t-2}$	0.077	0.00		
$\Delta s_{i,t-3}$	0.104	0.00		
$\Delta s_{i,t-4}$	-0.157	0.00		

exchange rates of the countries under observation are largely driven by common shocks. From this point of view, a coordinated exchange rate policy does not seem to be devious although the identification of shocks remains notoriously difficult. It seems reasonable to argue that shocks which stem from the US economy are fairly important in this context. The established long-run relationship points out that the common directions of monetary policies and business cycles relative to the US are an important determinant of exchange rates *vis-à-vis* the dollar. We feel encouraged by the fact that money supply and income turn out to be significant and enter with a sign which is in line with the framework of the monetary exchange rate model. Altogether, we conclude that the monetary approach is valid as a long-run anchor for the nominal exchange rate.

A major task for future research is the identification of break-points within our framework. The question whether potential instabilities coincide with major economic events for example with currency crisis is of specific importance in terms of policy implications.

References

Abuaf, N. and P. Jorion (1990). Purchasing power parity in the long run. *Journal of Finance 45*(1), 157–174.

Arellano, M. and S. Bond (1991). Some tests of specification for panel data: Monte Carlo evidence and an application to employment equations. *Review of Economic Studies 58*(2), 277–297.

Bai, J. and S. Ng (2004). A PANIC attack on unit roots and cointegration. *Econometrica 72*, 1127–1177.

Banerjee, A., M. Marcellino, and C. Osbat (2004). Some cautions on the use of panel methods for integrated series of macroeconomic data. *Econometrics Journal 7*(2), 322–340.

Basher, S. A. and J. Westerlund (2009). Panel cointegration and the monetary exchange rate model. *Economic Modelling 26*(2), 506–513.

Breitung, J. and M. H. Pesaran (2008). Unit roots and cointegration in panels. In L. Matyas and P. Sevestre (Eds.), *The Econometrics of Panel Data: Fundamentals and Recent Developments in Theory and Practice*, pp. 279–322. Dordrecht: Kluwer Academic Publishers.

Campbell, J. Y. and P. Perron (1991). Pitfalls and opportunities: What macroeconomists should know about unit roots. In O. J. Blanchard and S. Fisher (Eds.), *NBER Macroeconomics Annual*, Volume 6, pp. 141–220. Cambridge: MIT Press.

Cerra, V. and S. C. Saxena (2010). The monetary model strikes back: Evidence from the world. *Journal of International Economics 81*(2), 184–196.

Choi, I. (2001). Unit root tests for panel data. *Journal of International Money and Finance 20*(2), 249–272.

Dickey, D. A. and W. A. Fuller (1979). Distribution of the estimators for autoregressive time series with a unit root. *Journal of the American Statistical Association 74*(366), 427–431.

Dreger, C. (2010). Does the nominal exchange rate regime affect the real interest parity condition? *North American Journal of Economics and Finance 21*, 274–285.

Frankel, J. A. (1979). On the mark: A theory of floating exchange rates based on real interest differentials. *American Economic Review 69*(4), 610–22.

Frömmel, M., R. MacDonald, and L. Menkhoff (2005a). Do fundamentals matter for the d-mark/euro-dollar? A regime switching approach. *Global Finance Journal 15*(3), 321–335.

Frömmel, M., R. MacDonald, and L. Menkhoff (2005b). Markov switching regimes in a monetary exchange rate model. *Economic Modelling 22*(3), 485–502.

Gengenbach, C., F. C. Palm, and J.-P. Urbain (2006). Cointegration testing in panels with common factors. *Oxford Bulletin of Economics and Statistics 68*, 683–719. Supplement.

Goldberg, M. D. and R. Frydman (2001). Macroeconomic fundamentals and the DM/$ exchange rate: Temporal instability and the monetary model. *International Journal of Finance & Economics 6*(4), 421–35.

Granger, C. W. J. (1986). Developments in the study of cointegrated economic variables. *Oxford Bulletin of Economics and Statistics 48*(3), 213–28.

Groen, J. J. J. (1999). Long horizon predictability of exchange rates: Is it for real? *Empirical Economics 24*(3), 451–469.

Groen, J. J. J. (2002). Cointegration and the monetary exchange rate model revisited. *Oxford Bulletin of Economics and Statistics 64*(4), 361–80.

Hakkio, C. S. (1984). A re-examination of purchasing power parity: A multi-country and multi-period study. *Journal of International Economics 17*(3-4), 265–277.

Harris, R. I. D. and R. Sollis (2003). *Applied time series modelling and forecasting*. Chichester: J. Wiley.

Haug, A. A., J. G. MacKinnon, and L. Michelis (2000). European monetary union: A cointegration analysis. *Journal of International Money and Finance 19*(3), 419–432.

Husted, S. and R. MacDonald (1998). Monetary-based models of the exchange rate: A panel perspective. *Journal of International Financial Markets, Institutions and Money 8*(1), 1–19.

Johansen, S. (1995). *Likelihood-based inference in cointegrated vector autoregressive models*. Oxford: Oxford University Press.

Juselius, K. (2007). *The Cointegrated VAR model: Methodology and applications* (2 ed.). New York: Oxford University Press.

Kühl, M. (2008). Strong comovements of exchange rates: Theoretical and empirical cases when currencies become the same asset. *cege - Center for European, Governance and Economic Development Research Discussion Paper No. 76*.

Kühl, M. (2010). Bivariate cointegration of major exchange rates, cross-market efficiency and the introduction of the euro. *Journal of Economics and Business 62*(1), 1–19.

Levin, A., C. Lin, and C. J. Chu (2002). Unit root tests in panel data: Asymptotic and finite-sample properties. *Journal of Econometrics 108*(1), 1–24.

MacKinnon, J. G. (1996). Numerical distribution functions for unit root and cointegration tests. *Journal of Applied Econometrics 11*(6), 601–18.

MacKinnon, J. G., A. A. Haug, and L. Michelis (1999). Numerical distribution functions of likelihood ratio tests for cointegration. *Journal of Applied Econometrics 14*(5), 563–577.

Maddala, G. S. and S. Wu (1999). A comparative study of unit root tests with panel data and a new simple test. *Oxford Bulletin of Economics and Statistics 61*, 631–52.

Mark, N. C. and D. Sul (2001). Nominal exchange rates and monetary fundamentals: Evidence from a small post-bretton woods panel. *Journal of International Economics 53*(1), 29–52.

Mark, N. C. and D. Sul (2003). Cointegration vector estimation by panel DOLS and long-run money demand. *Oxford Bulletin of Economics and Statistics 65*(5), 655–680.

Meese, R. A. and K. Rogoff (1983). Empirical exchange rate models of the seventies: Do they fit out of sample? *Journal of International Economics 14*(1-2), 3–24.

Ng, S. and P. Perron (2001). Lag length selection and the construction of unit root tests with good size and power. *Econometrica 69*(6), 1519–1554.

Nickell, S. J. (1981). Biases in dynamic models with fixed effects. *Applied Economics 49*(6), 1417–1426.

Pedroni, P. (1999). Critical values for cointegration tests in heterogeneous panels with multiple regressors. *Oxford Bulletin of Economics and Statistics 61*, 653–670.

Pedroni, P. (2004). Panel cointegration, asymptotic and finite sample properties of pooled time series tests with an application to the PPP hypothesis. *Econometric Theory 20*, 597–625.

Phengpis, C. and V. Nguyen (2009). Policy coordination and risk premium in foreign exchange markets for major EU currencies. *Journal of International Financial Markets, Institutions and Money 19*(1), 47–62.

Phillips, P. C. B. and P. Perron (1988). Testing for a unit root in time series regression. *Biometrika 75*(2), 335–346.

Rapach, D. E. and M. E. Wohar (2004). Testing the monetary model of exchange rate determination: A closer look at panels. *Journal of International Money and Finance 23*(6), 867–895.

Reimers, H. (1992). Comparisons of tests for multivariate cointegration. *Statistical Papers 33*(1), 335–359.

Reinsel, G. C. and S. K. Ahn (1992). Vector autoregressive models with unit roots and reduced rank structure: Estimation, likelihood ratio test, and forecasting. *Journal of Time Series Analysis 13*(4), 353–375.

Sarno, L. and M. P. Taylor (2003). *The Economics of Exchange Rates*. Cambridge: Cambridge University Press.

Taylor, M. P. and D. A. Peel (2000). Nonlinear adjustment, long-run equilibrium and exchange rate fundamentals. *Journal of International Money and Finance 19*(1), 33–53.

Taylor, M. P., D. A. Peel, and L. Sarno (2001). Nonlinear mean-reversion in real exchange rates: Toward a solution to the purchasing power parity puzzles. *International Economic Review 42*(4), 1015–42.

Urbain, J. and J. Westerlund (2006). Spurious regression in nonstationary panels with Cross-Unit cointegration. *METEOR Research Memoranda No. 057*.

Wu, Y. (1996). Are real exchange rates nonstationary? Evidence from a panel-data test. *Journal of Money, Credit and Banking 28*(1), 54–63.

Yuan, C. (2011). The exchange rate and macroeconomic determinants: Time-varying transitional dynamics. *The North American Journal of Economics and Finance*. In Press.