Is the Monetary Model Useful in Explaining Exchange Rates? - Panel Cointegration Evidence

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A number of studies have sought to provide a reasonable explanation for exchange rate determination. The most frequently used approach is based on monetary models. However, it is difficult to find a cointegration relationship between exchange rates and relative differentials of money and income using this approach. This does not mean that a cointegration relationship does not exist. Conventional single equation approaches simply have a low performance power. We employed the panel cointegration approach to overcome this potential problem. We formulated a system of monetary models for 8 nations and found that cointegration relationships existed. Given these cointegration relationships, we estimated cointegrating vectors that are consistent with theoretical signs and magnitude.

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I. Introduction

Exchange rate determination has been studied through a variety of research avenues. Although many studies have looked for

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support for economic models, it has appeared that there are no systematic economic forces at work in determining exchange rates. In comparisons of the forecasting accuracy of monetary models and time series models, the random walk model outperformed the monetary approach and all other time series models (Meese and Rogoff 1983; and Frankel 1979).

In most studies, the method typically adopted is a single equation cointegration approach. This is because most economic variables, including exchange rates, appear to be $I(1)$, and require that we establish a cointegration relationship between them. However, there have been many unresolved debates regarding the validity of the monetary model in explaining exchange rate determination. In fact, the cointegration approach rarely supports the monetary model.

This lack of support for the monetary model may be due to the unreliability of the statistical procedures employed. It is commonly accepted that test statistics, such as ADF and Phillips-Perron tests for the null of unit roots and no-cointegration, have low finite sample power. Shiller and Perron (1985) demonstrated that the power of these tests depends on the range of data and not the frequency. Therefore, it is possible to conjecture that the low power performance may be due to the limited number of time series observations.

Virtually all testing procedures assume an asymptotic distribution, and hence additional observations are helpful for statistical inference. Fortunately, there is a straightforward way of increasing time series observations when the time span is limited. Since there are many cross-sections with meaningful time spans, it is easy to formulate panel data by collecting the same economic time series across as many cross-sections as possible.

In fact, panel data has been used in many areas. In literature related to growth, there are many studies on the variation in output and productivity across countries (Barro 1991), across states (Barro and Sala-i-Martin 1992), and across industries (Caballero and Lyons 1989). Oh *et al.* (1999) examined saving-investment cointegration for 7 countries using panel unit roots tests. Mark and Sul (2001) reported money-exchange rate cointegration relationships for 18 countries.

There are many studies that suggest that pooling time series could improve the finite sample properties of asymptotic tests.

Levin and Lin (1992) suggested panel-based unit root tests and demonstrated significant power improvement in finite samples. Oh (1996) also confirmed the power improvement, and provided evidence in favor of the PPP hypothesis. Pedroni (1995) suggested cointegration tests that are applicable to panel data. These studies generally show that pooling time series improves the power performance of conventional single series tests and generates better results in favor of economic theory.

By pooling time series, one can apply asymptotic tests even though the time span is short, as long as one can formulate relatively large cross-sections. Conceptually, it is straightforward to construct panel data by collecting data from hundreds of countries, regions, or other sources.

Furthermore, one can surmise that identical economic relationships exist across cross-sections, provided that cross-section-specific components are controlled. For example, Solow's growth model does not assume a specific country, and is applicable to many different countries. Given growth factors, such as population and saving rates, the economies of different countries with identical production functions might ultimately converge on an identical steady state. Evans (1997) and Evans and Karras (1996) examined panel data and showed that convergence does occur according to neo-classical growth theory.

However, there are few system approaches for exchange rate determination. Considering that a variety of panels across countries can easily be formed, it is rather surprising that the panel cointegration approach has not been employed to explain exchange rate determination. Recently, Mark and Sul (2001) examined a panel of 18 countries and found evidence in favor of a cointegration relationship between nominal exchange rates and monetary fundamentals. Given the cointegration relationship, they demonstrated that monetary fundamentals do have forecasting ability under many different specifications.

In this study, we employ the panel approach suggested by Pedroni (1995, 1996) to see whether a cointegration relationship exists among exchange rates, money differential, and income differential, using a simple monetary model. Furthermore, when a cointegration relationship appears to exist, we estimate the cointegration relationship directly by panel FMOLS. We present models and statistical procedures in section II, empirical results in section III, and concluding remarks in section IV.

II. Model and Statistical Procedures

A simple monetary model for a single equation based on the PPP hypothesis and demand for money could be expressed as:

$$s_t = \mu + \alpha(m_t - m_t^*) + \beta(y_t - y_t^*) + e_t, \tag{1}$$

where s, m, and y denote the exchange rate, money differential, and income differential, respectively. The superscript $*$ denotes foreign variables. The coefficient α is positive and β is negative. In particular, α equals one under the neutrality of money (Florentis *et al.* 1994).

In view of the potential weakness of a conventional single equation approach, we construct the following system of equations:

$$\begin{aligned} s_{1,t} &= \mu_1 + \alpha(m_{1,t} - m_t^*) + \beta(y_{1,t} - y_t^*) + e_{1,t} \\ s_{2,t} &= \mu_2 + \alpha(m_{2,t} - m_t^*) + \beta(y_{2,t} - y_t^*) + e_{2,t} \\ &\vdots \\ s_{n,t} &= \mu_n + \alpha(m_{n,t} - m_t^*) + \beta(y_{n,t} - y_t^*) + e_{n,t}\ , \end{aligned} \tag{2}$$

where μ_i denotes country-specific effects and $s_{i,t}$, $m_{i,t}$, and $y_{i,t}$ denote variables in country i at time t. As usual, $*$ denotes foreign variables. Since most economic time series appear to be $I(1)$, we assume that all variables are non-stationary.

The model describes an n-country system. All the coefficients maintain the same magnitude, except for country-specific effects. It is reasonable to assume an identical economic relationship, given a highly integrated financial system. For example, there is no *a priori* requirement that one nation has an entirely different economic relationship from the others in a certain group of countries.

The constant term captures the fixed effects of each cross-section. This is particularly important when one examines a cointegration relationship. Even if each country maintains an identical economic relationship, there might be country-specific effects due to technology, industry structure, a central banking system, and so on. We expect the fixed effects to control heterogeneity across borders. We could, of course, allow all the coefficients to vary over the cross-section, but there would then be no additional benefits from pooling the time series.

Compared with the single equation approach, a system approach has many benefits. First, by pooling the time series, we can improve finite sample performance. Levin and Lin (1992) showed that the power of unit root tests was significantly improved compared to univariate testing procedures. Oh (1996) also demonstrated that this is true, and applied it to find support for the PPP hypothesis. Pedroni (1995) suggested panel cointegration tests and found evidence in favor of the PPP hypothesis.

Second, as has been widely noted, for valid statistical inferences we need to remove nuisance parameters asymptotically. This can be done by parametric or non-parametric methods. Many estimation techniques are available for the single equation model, such as FMOLS (Phillips and Hansen 1990), lead-lag augmented OLS (Stock and Watson 1993) and CCR (Park 1992). With modification, these techniques are also applicable to a pooled model.

Third, econometrically, non-stationary time series offer a way to avoid problems such as the low power performance of conventional testing procedures. The poor performance of unit root tests such as ADF and PP is widely acknowledged when the time series have a limited span.

Fourth, pooling time series can extract useful information inherent in the economic nature of the system. In fact, there is no *a priori* requirement to either assume different economic relationships or to maintain identical relationships across cross-sections. However, unless there is some restriction across the cross-section there is no merit in pooling time series.

In this study, we adopted the panel cointegration tests and panel FMOLS of Pedroni (1996). Brief descriptions follow. Assuming that identical cointegration relationships exist, consider the following OLS estimator:

$$\hat{\theta} = \left[\sum_{i=1}^{N}\sum_{t=1}^{T} \bar{x}_{it}\, \bar{x}_{it}' \right]^{-1} \left[\sum_{i=1}^{N}\sum_{t=1}^{T} \bar{x}_{it}\, \bar{y}_{it}' \right], \tag{3}$$

where $\hat{\theta} = (\alpha,\ \beta)'$ and $\bar{x}_{it} = x_{it} - \bar{x}_i$.

Non-stationarity generally involves nuisance parameters that must be removed asymptotically. As in Pedroni (1996), it is possible to apply the FMOLS principles suggested by Phillips and Hansen (1990) to remove the nuisance parameters. The estimators are given by:

$$\hat{\theta} = \left[\sum_{i=1}^{N}\sum_{t=1}^{T} L_{22i}^{-1}\bar{x}_{it}\,\bar{x}_{it}'L_{22i}^{-1\prime} \right]^{-1}\left[\sum_{i=1}^{N}\sum_{t=1}^{T} L_{22i}^{-1}\{\bar{x}_{it}\,\bar{y}_{it}^{+} - T\gamma_i\}L_{11i}^{-1} \right], \tag{4}$$

where $\bar{y}_{it}^{+} = (y_{it} - \bar{y}_i) - L_{21i}'L_{22i}^{-1}\Delta x_{it}$ and $L_iL_i' = \Omega_i$. L_{pqi} is partition of L_i and Ω_i is the longrun covariance matrix of $(\Delta y_{it}\ \Delta x_{it}')'$.

As N and T go to infinity, the asymptotic distribution of the estimator approaches

$$\sqrt{NT}(\hat{\theta} - \theta) \Rightarrow N(0,\ M^{-1}), \tag{5}$$

where $M = E\{\int(W - \int W)(W - \int W)'\}$ and W denotes vector Brownian motion.

The results are quite interesting. The coefficient estimator converges at a much faster rate than the single equation estimators, and follows a normal distribution with proper normalization. It is therefore straightforward to test a parameter value using traditional testing methods such as t-statistics and Wald tests.

Given that all of the time series under study are non-stationary, we need to know whether the system is in fact cointegrated. In order to test the null hypothesis that the system is not cointegrated, we use OLS residuals and construct various test statistics, as suggested by Pedroni (1995). There are two different specifications: one is for the homogeneous model and the other is for heterogeneous models. First, in the homogeneous model, all the coefficients are identical over a cross-section, whereas the heterogeneous model does not impose parameter restrictions. Here, we modify the homogeneous model slightly, and allow constants to differ over a cross-section. This does not alter the asymptotic properties, since the estimated residuals are equivalent to row residuals as N and T go to infinity. From the panel autoregression, $e_{it} = \rho e_{it-1} + \eta_{it}$, the coefficient estimator is given by:

$$(\hat{\rho} - 1) = \left(\sum_{n=1}^{N}\sum_{t=1}^{T} e_{it-1}^{2} \right)^{-1} \sum_{n=1}^{N}\sum_{t=1}^{T} (e_{it-1}\Delta e_{it} - \hat{\lambda}_i), \tag{6}$$

where $\hat{\lambda}_i$ is the usual correction term as in Phillips-Perron tests.

Then the following asymptotic properties hold(Corr. 3.1 of Pedroni 1995):

(a) $T\sqrt{N}(\hat{\rho}-1) \Rightarrow N(0,\ 2)$ as $N,\ T \to \infty$ for general $\Omega_i = \Omega > 0$ disturbances.
(b) $\sqrt{NT(T-1)}(\hat{\rho}-1) \Rightarrow N(0,\ 2)$ as $N \to \infty$ for *iid* disturbances, regardless of T.
(c) $t_\rho \Rightarrow N(0,\ 1)$ as $N \to \infty$ for *iid* disturbances, regardless of T, and as for general $\Omega_i = \Omega > 0$ disturbances.

Next, for the heterogeneous panel, the following test statistics are suggested:

$$Z_{\hat{\nu}} = \left(\sum_{n=1}^{N} \sum_{t=1}^{T} \hat{L}_{11i}^{-2} \hat{e}_{it-1}^{2} \right)^{-1}$$

$$Z_{\hat{\rho}} = \left(\sum_{n=1}^{N} \sum_{t=1}^{T} \hat{L}_{11i}^{-2} \hat{e}_{it-1}^{2} \right)^{-1} \sum_{n=1}^{N} \sum_{t=1}^{T} \hat{L}_{11i}^{-2} \left(\hat{e}_{it-1} \hat{e}_{it} - \hat{\lambda}_i \right)$$

$$Z_{t_{\hat{\rho}}} = \left(\hat{\sigma}^2 \sum_{n=1}^{N} \sum_{t=1}^{T} \hat{L}_{11i}^{-2} \hat{e}_{it-1}^{2} \right)^{-\frac{1}{2}} \sum_{n=1}^{N} \sum_{t=1}^{T} \hat{L}_{11i}^{-2} \left(\hat{e}_{it-1} \Delta\hat{e}_{it} - \hat{\lambda}_i \right)$$

Asymptotic results for the model with intercepts under the null of no-cointegration are given as (Pedroni 1996):

$$TN^{3/2} Z_{\hat{\nu}} - 8.62\sqrt{N} \Rightarrow N(0,\ 60.75)$$

$$TN^{1/2} Z_{\hat{\rho}} - 6.02\sqrt{N} \Rightarrow N(0,\ 31.27)$$

$$Z_{t_{\hat{\rho}}} - 1.73\sqrt{N} \Rightarrow N(0,\ 0.93)$$

Here we assume the coefficient ρ is constant and test H_0: $\rho = 1$ against H_A: $|\rho| < 1$. It may be possible to test more general hypothesis such as H_0: $\rho_i = 1$ for all i against H_A: $|\rho_i| < 1$ for some i to reflect panel cointegration. Not identical but similar cases were examined in Choi (2000) and Im, Pesaran and Shin (2001). Considering our object to enjoy benefits of large number of observations, we adapted constant panel autocorrelation coefficient for the study.

III. Data and Results

In this study, we considered Canada, France, Germany, Japan, Italy, Switzerland, the UK, and the U.S. Quarterly series were abstracted from the IFS CD-ROM of December 1997. The series run from 1973:1-1997:2, although some values were missing and were replaced by the PWT data set. Annual series were constructed from the sum or average of the quarterly series. We calculated 1997 values by assuming that the 3^{rd} and 4^{th} quarters were the same as earlier quarters. The U.S. was considered a foreign country in all the calculations, hence, $s_{i,t}$ denotes U.S. dollar denominated exchange rates. Logarithmic values were used for all the series. All calculations were made using GAUSS and Coint.

In order to see whether there were unit roots in our time series, we applied Dicky-Fuller ADF and Phillips-Perron Z_α and Z_t to the quarterly and annual series. Z_α never rejected the null of unit roots, while Z_t and ADF occasionally rejected the nulls for monetary aggregates and CPI, respectively. Varying lag lengths did not alter the results significantly. As is commonly accepted in many studies, we concluded that GNP, exchange rates, monetary aggregates, and price levels are integrated. For differential series, the unit root hypotheses were seldom rejected. Considering most empirical works and test results, we assumed that all variables are integrated.[1]

First, we tested for the null of no-cointegration and the results are reported in Table 1. Part A displays the cointegration test results for individual countries. Conventional single equation tests rarely reject the null of no-cointegration, which is consistent with earlier studies. This does not, however, demonstrate that the monetary model of exchange rate determination is of no use.[2] As has been argued by many others, the result may be due to the tests' lack of power.[3]

[1]Results are not reported to save space. They are available upon request.

[2]There are studies that support cointegration relationships. For example, Florentis *et al.* (1994) failed to support cointegration relationships for the Canadian dollar, whereas McNown and Wallace (1994) found evidence in favor of cointegration for high inflation economies using a slightly modified monetary model, but obtained parameter estimates with contradictory signs in many cases.

[3]Structural breaks and limited span of the time series are usually mentioned as sources of the low power of these tests.

TABLE 1

COINTEGRATION TESTS FOR INDIVIDUAL AND PANEL MONETARY MODEL

Part A. Testing for Cointegration for Individual Country						
	Quarterly			Annual		
	Z_α	Z_t	ADF	Z_α	Z_t	ADF
Canada	-6.599	-1.823	-2.116	-12.217	-2.524	-2.506
France	-11.356	-2.479	-3.291	-15.018	-2.867	-2.554
Germany	-9.427	-2.358	-2.406	-10.076	-2.365	-1.611
Italy	-9.844	-2.242	-2.785	-14.309	-2.796	-2.959
Japan	-4.62	-1.229	-1.872	-12.492	-2.455	-1.925
Switzerland	-11.065	-2.363	-2.673	-12.762	-2.611	-5.421
U.K.	-13.058	-2.605	-2.769	-17.249	-3.057	-1.844
CV 95%	-24.523	-3.838	-3.838	-24.523	-3.838	-3.838

Part B. Testing for Panel Cointegration				
Homogeneous Panel	$T\sqrt{N}(\hat{\rho}-1)$	-11.148	$T\sqrt{N}(\hat{\rho}-1)$	-10.446
	$\sqrt{NT(T-1)}(\hat{\rho}-1)$	-11.091	$\sqrt{NT(T-1)}(\hat{\rho}-1)$	-10.235
	$t_{\hat{\rho}}$	-3.537	$t_{\hat{\rho}}$	-3.298
Heterogeneous Panel	$Z_{\hat{\nu}}$	-22.437	$Z_{\hat{\nu}}$	-21.286
	$Z_{\hat{\rho}}$	-35.361	$Z_{\hat{\rho}}$	-34.203
	$Z_{t_{\hat{\rho}}}$	-7236.36	$Z_{t_{\hat{\rho}}}$	-594.462

Notes: (1) The null of no-cointegration is rejected for all cases in Part B.
(2) Parzen kernel with lag length 4 was used in computation.

In order to see whether the conventional results are indeed valid, we applied panel cointegration tests. When we applied panel cointegration tests to the system, the null of no-cointegration was strongly rejected for both homogeneous and heterogeneous models. Part B reports the results of the panel cointegration tests. The upper section displays the results for homogeneous panels, and the lower section the results for heterogeneous panels. The null was rejected in all cases. Considering the evidences provided by Levin and Lin (1992) and Pedroni (1995), to name a few, the results can be seen as favoring a cointegration relationship. Unlike in previous studies, the monetary model appears to have strong empirical support once the time series are pooled. The results are identical with those of Mark and Sul (2001) in the sense that cointegration relationships for single equations were not supported, but the system exhibited cointegration relationships for panel data.

Next, we proceeded to estimate equation (2) based on FMOLS

TABLE 2

FMOLS ESTIMATES FOR INDIVIDUAL EQUATION AND POOLING TIME SERIES

Part A. Individual FMOLS Estimates (reference only)

Country	Quarterly		Yearly	
	m-m*	y-y*	m-m*	y-y*
Canada	0.131	1.668	0.093	2.155
France	0.732	-1.381	0.728	-1.233
Germany	-0.469	-1.134	-0.737	-0.672
Italy	0.843	-1.613	0.773	-0.991
Japan	1.085	-3.543	0.558	-5.065
Switzerland	0.833	0.266	0.883	0.337
U.K.	0.031	-3.057	0.008	-3.171

Part B. Panel FMOLS Estimates

Lag	m-m*	y-y*	m-m*	y-y*
1	0.372(69.715)	-0.804(-70.517)	0.518(13.525)	-0.671(-17.464)
2	0.377(56.177)	-0.779(-66.216)	0.562(11.954)	-0.685(-19.818)
3	0.388(47.129)	-0.761(-68.798)	0.539(10.626)	-0.703(-16.284)
4	0.396(42.486)	-0.758(-71.838)	0.432(11.211)	-0.882(-22.458)
5	0.401(39.715)	-0.758(-76.093)	0.475(11.996)	-0.890(-39.165)

Note: Numbers in parentheses denote *t*-ratio, which follow N(0, 1) asymptotically.

principles. The results are reported in Table 2. Part A gives estimates from single equation FMOLS for comparison.[4] The parameter estimates for Canada, Germany, and Switzerland carry a theoretically incorrect sign, and the magnitudes from each cross-section are also very different. The results are consistent with McNown and Wallas (1994), who reported cointegration relationships for three high-inflation economies, but in many cases obtained parameter estimates with contradictory signs.

Part B shows the panel FMOLS estimates. Lag lengths were allowed to vary from 1 to 5. Nominal exchange rates were regressed on money differentials and GDP differentials. All the parameter estimates were statistically significant, but the neutrality of money was rejected. The estimates of the quarterly model were similar to those of Boothe (1983), but larger than those of Florentis *et al.* (1994) for the case of Canada.[5]

[4]Single equation FMOLS are not valid, since cointegration was not favored by conventional tests.

All the parameter estimates were statistically significant and carried the correct sign. Varying the lag length from 1 to 5 did not alter the results. This was true for both the quarterly and annual series. The results are quite interesting, since the monetary model actually has meaningful explanatory power for exchange rate determination. The money differential elasticity of exchange rates is between 0.37~0.4 and 0.48~0.52 for quarterly and annual series, respectively. In addition, the income elasticity of exchange rates ranges between 0.76~0.80 and 0.67~0.89 for quarterly and annual series, respectively. Income elasticity is almost twice money elasticity for quarterly series, and is 50% higher for annual series. When quarterly money differentials increase by 1%, exchange rates depreciate about 0.37~0.4%. When quarterly GDP differentials increase by 1%, exchange rates appreciate by about 0.76~0.80%. Income differentials appear to be much more important than money differentials.

IV. Concluding Remarks

In this study, we tested whether the monetary model is useful in explaining exchange rate determination. By adopting Pedroni (1995, 1996), we found evidence in favor of cointegration relationships, from a simple monetary model. We also estimated the coefficients and found that the money differential elasticity was between 0.37~0.4 and 0.48~0.52 for quarterly and annual series, respectively. In addition, the income elasticity of exchange rates ranged between 0.76~0.8 and 0.67~0.89 for quarterly and annual series, respectively.

The results suggest that monetary models are useful in explaining exchange rate behavior. Unlike conventional single equation cointegration approaches, we found evidence that pooling time series may reveal the usefulness of monetary models. In this regard, our findings may be due to the power improvement from time series pooling as demonstrated in Levin and Lin (1992), Oh (1996), and Pedroni (1995). Additionally, the results are consistent with Mark and Sul (2001) who examined exchange rate-monetary

[5]The results are not directly compatible because they did not consider nonstationarity of the time series.

fundamentals link within a panel data framework and rejected the null of no cointegration.

Some reservation may be necessary. Firstly, one may devise cointegration tests for more general hypothesis. For example, one may test the null of no cointegration for all cross sections against the alternative of cointegration for some individual sectors. Secondly, new tools to incorporate contemporaneous correlation over the cross section need to be developed. As suggested in Bai (2001), Bai and Ng (2000), and Banerjee (1999), contemporaneous correlation over the cross section would be an important issue on which further studies are necessary in non-stationary panel data.

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