Testing the Monetary Model of Exchange Rate Determination: A Closer Look at Panels

David E. Rapach
Seattle University
rapachd@seattleu.edu

Mark E. Wohar*
University of Nebraska-Omaha
wohar@unomaha.edu

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Abstract

In this paper, we undertake an extensive evaluation of panel tests of the long-run monetary model of exchange rate determination. We first show how poorly the monetary model performs on a country-by-country basis for U.S. dollar exchange rates over the post-Bretton Woods period for a large number of industrialized countries. In sharp contrast, we find considerable support for the monetary model using panel procedures, as in Groen (2000) and Mark and Sul (2001). Given the disparity in the country-by-country and panel approaches, we carefully analyze the homogeneity restrictions inherent in the panel procedures. The evidence on the appropriateness of the homogeneity restrictions is mixed. In the end, whether the monetary model conforms to post-Bretton Woods data largely depends on one's prior beliefs.

Key words: Monetary model; Cointegration; Panel tests; Homogeneity restrictions

JEL classification: C32, C33, F31, F41

*Corresponding author. Address correspondence to: Department of Economics, University of Nebraska-Omaha, CBA 512K, Omaha, NE 68182-0286; Phone: 402-554-3712; Fax: 402-554-2853. We thank Nelson Mark and Donggyu Sul for generously providing the data used in this paper and two anonymous referees for helpful comments on an earlier draft. The usual disclaimer applies. The results reported in this paper were generated using GAUSS 3.5.

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1 Introduction

Empirical studies investigating the monetary model of exchange rate determination on a country-by-country basis over the modern floating exchange rate period find virtually no support for this long-standing theoretical model. In particular, empirical studies typically fail to find evidence of a cointegrating relationship between the nominal exchange rate and a simple set of monetary fundamentals—as required by the "long-run" monetary model—on a country-by-country basis during the modern float; see, for example, Meese (1986), Baillie and Selover (1987), McNown and Wallace (1989), Baillie and Pecchenino (1991), and Sarantis (1994). Even studies that find evidence of cointegration between nominal exchange rates and monetary fundamentals during the modern float lend little support to the monetary model, as the estimated cointegrating vectors themselves often fail to conform to the monetary model. For example, Cushman (2000) finds evidence of cointegration between the U.S. dollar-Canadian dollar exchange rate and a set of monetary fundamentals over the modern float, but the estimated cointegrating coefficients differ widely from those predicted by the monetary model. Cushman (2000) therefore concludes that there is no support for the monetary model in U.S.-Canadian data.[1]

It is not surprising that the monetary model performs poorly on a country-by-country basis during the modern float, as long-run purchasing power parity (PPP)—a building block of the monetary model—also fares poorly on a country-by-country basis during this period. That is, there is little evidence that nominal exchange rates and relative price levels cointegrate, as required by long-run PPP, on a country-by-country basis.[2] The failure of long-run PPP is typically attributed to the short spans of data available during the post-Bretton Woods float, so that standard tests (which take no cointegration as the null hypothesis) have extremely low power to reject the null hypothesis of no cointegration using data from the post-Bretton Woods

[1]As another example, MacDonald and Taylor (1994) find evidence of cointegration between the U.S. dollar-U.K. pound exchange rate and a set of monetary fundamentals using data covering 1976-1990, but their cointegrating vector is difficult to interpret theoretically. MacDonald and Taylor (1994) are only able to make the relatively weak assertion that the cointegrating vector "does not, in fact, do great violence to the monetary model."

[2]For surveys of the PPP literature, see Breuer (1994), Froot and Rogoff (1995), Taylor (1995), Rogoff (1996), and Sarno and Taylor (2002)

period.[3] A popular response to the problem of low power in the PPP literature is the use of panel data, and, as initially shown by Levin and Lin (1992), combining cross-sectional and time-series information in the form of a panel can greatly increase the power of unit root and cointegration tests. In fact, a number of panel studies find support for long-run PPP using data from the post-Bretton Woods period, including Pedroni (1995), Frankel and Rose (1996), Oh (1996), Wu (1996), Papell (1997), and Taylor and Sarno (1998).

Two recent studies by Groen (2000) and Mark and Sul (2001) follow the PPP literature and test the monetary model using panels of post-Bretton Woods data. Groen (2000) considers a panel of U.S. dollar nominal exchange rate, relative money supply, and relative real output level data for 14 industrialized countries covering the period 1973:1–1994:4. He obtains panel cointegration coefficient estimates that are reasonably consistent with the monetary model for his full panel and three subpanels (G10, G7, and EMS). In addition, his panel cointegration test indicates that nominal exchange rates are cointegrated with relative money supplies and relative real output levels for his full panel and a G10 subpanel. He finds even more support for the monetary model when he uses the German mark as the numeraire currency. Mark and Sul (2001) employ a panel of U.S. dollar nominal exchange rate, relative money supply, and relative real output data for 18 countries spanning 1973:1–1997:1. They develop a panel cointegration test based on an error-correction specification that assumes pre-specified values for the cointegrating coefficients. Their panel cointegration test finds evidence of cointegration among U.S. dollar exchange rates, relative money supplies, and relative income levels for their full panel of 18 countries. They also find evidence of cointegration using the Swiss franc or Japanese yen as the numeraire currency.

The Groen (2000) and Mark and Sul (2001) panel results are noteworthy, given that the monetary model performs so poorly on a country-by-country basis during the post-Bretton Woods float, and their results suggest that the monetary model may be able to explain nominal exchange rate trends during the modern float after all. However, precisely because there is so little support for the monetary model on a country-by-country basis, it is important to scrutinize panel tests of the monetary model. To this end, we undertake an extensive econometric evaluation of panel tests of the monetary model in the present paper. We pay special

[3] As shown by Shiller and Perron (1985), Hakkio and Rush (1991), and Otero and Smith (2000), it is the span of the data, and not their frequency, that determines the power of unit root and cointegration tests.

attention to the issue of homogeneity, as both Groen (2000) and Mark and Sul (2001) assume homogeneous cointegrating coefficients across countries in their panel procedures. In fact, Mark and Sul (2001) assume pre-specified values for the homogeneous cointegrating coefficients. It is not immediately evident that these types of assumptions are supported by the data. The ultimate objective of our evaluation is to assess just how well the monetary model of U.S. dollar exchange rate determination conforms to post-Bretton Woods data.

We begin by providing the most extensive evidence to date on the poor performance of the monetary model on a country-by-country basis during the modern float. In order to find support for the monetary model, we require two pieces of evidence: (i) we need to obtain estimates of the cointegrating coefficients relating the nominal exchange rate to a set of monetary fundamentals that agree with the values predicted by the monetary model; (ii) we need to find evidence that the nominal exchange rate is in fact cointegrated with the monetary fundamentals. It is difficult to overstate how poorly the monetary model performs on both of these counts on a country-by-country basis for the countries in our sample. Using a variety of techniques, we obtain country-by-country estimates of cointegrating coefficients that typically diverge widely from the values predicted by the monetary model. In addition, standard tests provide little evidence of cointegration between nominal exchange rates and monetary fundamentals during the modern float on a country-by-country basis.

We next turn to panel tests of the monetary model that pool data across countries. Using a number of different estimation techniques, we obtain panel estimates of cointegrating coefficients that are often reasonably in line with the monetary model. Panel estimates for our full panel of 18 countries and four subpanels almost always have the correct sign, and their magnitudes are often plausible. The least squares dummy variable (LSDV) estimator that allows for a common time effect produces estimates that accord especially well with the monetary model. The contrast between our country-by-country and panel estimates of cointegrating coefficients is quite striking.

While our panel cointegrating coefficient estimates are often in line with the monetary model, it is also necessary to determine if a cointegrating relationship actually exists among nominal exchange rates, relative money supplies, and relative income levels by testing the stationarity of the residuals from the panel cointegrating regressions. We consider five residual-based panel cointegration tests: a Phillips and Perron (1988)-type test due to Pedroni (1995);

two Dickey and Fuller (1979)-type tests due to Kao (1999); the Groen (2000) feasible GLS test; and a Taylor and Sarno (1998) version of the Groen (2000) test. These tests differ in the way they handle cross-sectional dependence and serial correlation. It is important to consider a number of panel cointegration tests, as Groen (2000) and Mark and Sul (2001) each employ only a single panel cointegration test, thus making it difficult to judge the robustness of their cointegration inferences. Again in sharp contrast to the country-by-country results, we find considerable evidence for the existence of a cointegrating relationship among nominal exchange rates, relative money supplies, and relative income levels using our five panel cointegration tests. There is substantial agreement among the different tests, and all of the tests indicate cointegration for our full panel of 18 countries and three of our subpanels.

The stark contrast between our country-by-country and panel results compels us to examine the homogeneity assumptions inherent in panel procedures that rely on pooling, including those of Groen (2000) and Mark and Sul (2001). We therefore consider tests of the null hypothesis that the cointegrating slope coefficients are homogeneous across the 18 countries in our full panel, as well as across the countries in the various subpanels we consider. Interestingly, our tests typically reject the null hypothesis of homogeneity, so that the pooling of data inherent in panel procedures receives little support from the data. The dearth of support for homogeneity in the data leaves us with a conundrum: on the one hand, we would like to exploit the additional information that becomes available when we pool, and when we do this, we obtain cointegrating coefficient estimates that are much more plausible in economic terms than country-by-country estimates; on the other hand, formal statistical tests reject homogeneity, so we run the risk of obtaining spurious results by pooling the data. We consider the issue of pooling in detail, relating it to the recent work of Pesaran, Shin, and Smith (1999), Baltagi, Griffin, and Xiong (2000), and Coakley, Fuertes, and Smith (2001). This issue is likely to grow significantly in importance, given the increasing popularity of panel cointegration tests in applied work.

The rest of the paper is organized as follows: Section 2 reports our country-by-country test results of the monetary model; Section 3 reports our panel test results; Section 4 addresses the issue of pooling; Section 5 concludes.

2 Country-by-Country Results

2.1 The Theoretical Monetary Model

The basic form of the monetary model assumes stable money demand functions in the domestic and foreign countries, purchasing power parity, and uncovered interest parity. Given that the three assumptions at the core of the monetary model are unlikely to hold at each point in time, the monetary model should be viewed as a long-run or steady-state model of exchange rate determination. Suppressing a constant term, (1) shows the stable long-run relationship between the nominal exchange rates, relative money supplies, and relative income levels predicted by the monetary model:

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

where $\alpha > 0$, e_t is the nominal exchange rate measured in the number of units of foreign currency per unit of domestic currency, m_t is the nominal money supply, and y_t is real output (all at time t). Asterisks denote a foreign variable, and lower-case letters denote logarithms. The parameter α respresents the income elasticity of money demand in the domestic and foreign countries.[4] Given that e_t, $m_t^* - m_t$, and $y_t^* - y_t$ appear $I(1)$ over the post-Bretton Woods period in most countries, a test of (1) as a stable long-run equilibrium relationship is tantamount to testing for cointegration among $(e_t, m_t^* - m_t, y_t^* - y_t)'$ with cointegrating vector $(1, -1, \alpha)$. Groen (2000) and Mark and Sul (2001) test for cointegration among $(e_t, m_t^* - m_t, y_t^* - y_t)'$ in a panel framework.

2.2 The Data

We use the Mark and Sul (2001) data set. It contains quarterly data from 1973:1–1997:1 for nominal money supplies, industrial production, and nominal U.S. dollar exchange rates for 19 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. Mark and Sul (2001) note that the exchange rate experiences of these countries were dominated by floating during the sample period. Most of the data is from

[4] See, for example, Rapach and Wohar (2001) for a derivation of (1). Mark and Sul (2001) observe that a similar type of relationship can be derived from the Lucas (1982) and Obstfeld and Rogoff (1995) equilibrium models.

the International Monetary Fund's *International Financial Statistics* CD-ROM. The nominal money supply is the sum of money and quasi-money.[5] As real GDP is unavailable for a number of countries, industrial production is used to measure real income. See Mark and Sul (2001, Section 3) for more detail on the data set. With the U.S. serving as the domestic country in (1), we have e_t, $m_t^* - m_t$, and $y_t^* - y_t$ data for 18 countries.

2.3 Country-by-Country Cointegrating Coefficient Estimates

Consider testing the long-run monetary model for foreign country i with the United States serving as the domestic country. We begin with the levels regression for country i:

$$e_{it} = \beta_{i0} + \beta_{i1}(m_{it}^* - m_{it}) + \beta_{i2}(y_{it}^* - y_{it}) + u_{it}. \tag{2}$$

According to the monetary model, (i) $\beta_{i1} = 1$, $\beta_{i2} < 0$ and (ii) u_{it} is stationary. The first step in testing the monetary model is thus estimating β_{i1} and β_{i2} in (2). It is well-known that if (2) represents a true cointegrating relationship, OLS estimates of β_{i1} and β_{i2} are super-consistent. However, they suffer from asymptotic bias unless the regressors are strictly exogenous, so that the OLS standard errors cannot generally be used for valid inference. The Phillips and Hansen (1990) fully modified OLS (FM-OLS) and Saikkonen (1991) and Stock and Watson (1993) dynamic OLS (DOLS) estimators are designed to be free of endogeneity bias. They also provide consistent estimates of standard errors that can be used for inference.

Country-by-country OLS, FM-OLS, and DOLS estimates of the cointegrating slope coefficients are reported in Table 1. Columns (2) and (3) report the OLS estimates, while the FM-OLS and DOLS estimates are reported in columns (4) and (5) and columns (6) and (7), respectively. Following the applications in Hansen (1992), we use the quadratic kernel and the Andrews (1991) automatic bandwidth selector with Andrews and Monohan (1992) prewhitening when computing the FM-OLS estimates. Following Stock and Watson (1993), we include two leads and lags in the DOLS regression, and we use an autoregressive spectral density estimator of the long-run variance when calculating the standard errors.[6]

[5] As in Mark and Sul (2001), the nominal money supply data is seasonally adjusted by taking the average of the current and three previous quarters.

[6] We use the GAUSS program available from Bruce Hansen's home page (http://www.ssc.wisc.edu/~bhansen/) to calculate the FM-OLS estimates. Some of the code used to compute the DOLS estimates is based on the GAUSS program available from Mark Watson's home page (http://www.wws.princeton.edu/~mwatson).

The OLS, FM-OLS, and DOLS parameter estimates in Table 1 are often very much at odds with the values hypothesized by theory. In the majority of cases, the parameter estimates do not even have the correct sign. For the OLS estimates, only Australia, France, Italy, and Spain have β_{i1} and β_{i2} parameter estimates that are positive and negative, respectively, as required by theory. For the DOLS estimates, only France, Italy, and Spain have the correct signs. However, according to the DOLS standard errors, the β_{i2} estimate is insignificantly different from zero for these three countries, and the β_{i1} estimate is also insignificant for Italy and Spain. (As noted above, the OLS standard errors cannot be used for valid inference.) For the FM-OLS estimates, France and Spain again have parameter estimates with the correct signs, but again, the FM-OLS standards errors indicate that the β_{i2} estimate is insignificant for both countries and that the β_{i1} estimate is insignificant for Spain. Overall, there is no support for the monetary model for *any* country in Table 1.

We consider additional estimates of the cointegrating relationship implied by the monetary model in Table 2. Columns (2) and (3) report β_{i1} and β_{i2} parameter estimates based on the Johansen (1991) vector-error-correction maximum-likelihood (JOH-ML) approach, which is popular in applied work.[7] For the JOH-ML estimates, only France and Japan have parameter estimates with the correct signs. However, the β_{i2} estimate is insignificant for France, and the β_{i1} estimate is insignificant for Japan. Again, it is difficult to find support for the monetary model for any country using the JOH-ML estimator.

The second country-by-country estimator considered in Table 2 is the Mark, Ogaki, and Sul (2000) unrestricted dynamic SUR cointegration estimator. This is essentially an SUR version of the DOLS estimator, so it accounts for potential cross-sectional dependencies in the data. The dynamic SUR estimator clearly does not estimate the cointegrating coefficients for each country in isolation, but strictly speaking it is not a panel or pooled estimator, as the cointegrating coefficients are not restricted to be identical across countries. We consider this estimator in order to have a basis for testing the homogeneity restrictions inherent in panel estimators. Dynamic SUR estimation proceeds in two steps. In the first step, regressor endogeneity is purged through OLS estimation: for each country, the variables e_{it}, $m^*_{it} - m_{it}$,

[7]The Johansen (1991) multivariate procedure is not subject to endogeneity bias. We select the lag order for the vector-error-correction model using top-down testing and the Sims (1980) modified likelihood-ratio statistic. This is essentially a multivariate version of the top-down procedure recommended by Campbell and Perron (1991) and Ng and Perron (1995) for univariate unit root tests.

and $y_{it}^* - y_{it}$ are regressed in turn on lags and leads of $\Delta(m_{it}^* - m_{it})$ and $\Delta(y_{it}^* - y_{it})$ from all countries. The residuals are then stacked, with the e_{it} first-step residuals as regressands and the $m_{it}^* - m_{it}$ and $y_{it}^* - y_{it}$ first-step residuals as regressors, to form an SUR system. Feasible SUR estimates of β_{i1} and β_{i2} are then computed using an estimate of the long-run variance-covariance matrix for the SUR errors.[8] A drawback to the dynamic SUR estimator is that including lags and leads from all cross-sectional units in the first-step regressions quickly uses up degrees of freedom. In an application using forward and spot exchange rates, Mark, Ogaki, and Sul (2000) find that including leads only in the first-step regression produces results similar to those obtained when both lags and leads are included. In addition, they point out that it is more important to include leads when the regressand is an asset price. Given that we have 18 countries and 94 observations on e_{it}, $m_{it}^* - m_{it}$, and $y_{it}^* - y_{it}$ per country, we include one lead of $\Delta(m_{it}^* - m_{it})$ and $\Delta(y_{it}^* - y_{it})$ from each country in the first-step regression (which alone uses up 36 degrees of freedom). The dynamic SUR estimates of β_{i1} and β_{i2} are reported in columns (4) and (5) of Table 2. Only four of the 18 countries—Finland, Greece, Italy, and Spain—have the correct β_{i1} and β_{i2} signs using the dynamic SUR estimator. For Finland, both coefficient estimates are insignificant, while for Spain, the β_{i1} estimate is insignificant. Both coefficients are significantly different from zero for Greece and Italy, and the β_{i1} estimate is not significantly different from unity for both countries, as required by the monetary model. Overall, the dynamic SUR coefficient estimates are not much better in terms of the monetary model than the OLS, FM-OLS, DOLS, and JOH-ML estimates, although there is support for the monetary model for two countries (Greece and Italy) using the dynamic SUR estimator, while there is no support for any country using the other estimators.

Our final country-by-country estimator is the Pesaran and Shin (1999) autoregressive distributed lag (ARDL) estimator. Pesaran and Shin (1999) show that the coefficients of a cointegrating relationship can be estimated without asymptotic bias in an ARDL framework. In terms of the monetary model, one uses the following re-parameterized ARDL model for country i:

$$e_{it} = \phi_i e_{i,t-1} + \gamma_{i1}(m_{it}^* - m_{it}) + \gamma_{i2}(y_{it}^* - y_{it}) + \sum_{j=1}^{p-1} \lambda_{ij} \Delta e_{i,t-j} +$$

[8] The long-run variance-covariance matrix is estimated through a restricted VAR. See Mark, Ogaki, and Sul (2000, pp. 15-16) for details on the construction of the long-run variance-covariance matrix estimate.

$$\sum_{j=0}^{q_1-1} \delta_{i1j}\Delta(m^*_{i,t-j} - m_{i,t-j}) + \sum_{j=0}^{q_2-1} \delta_{i2j}\Delta(y^*_{i,j-j} - y_{i,t-j}) + \zeta_{it}, \quad (3)$$

where $\beta_{i1} = -\gamma_{i1}/\phi_i$ and $\beta_{i2} = -\gamma_{i2}/\phi_i$. Pesaran and Shin (1999) compare their ARDL estimator to the FM-OLS estimator, and they find that the ARDL estimator often performs better than the FM-OLS estimator in Monte Carlo simulations. In order to account for a measure of cross-sectional dependency, we allow for a common time effect across the 18 countries by subtracting out the mean across countries in each period for each of the variables e_{it}, $m^*_{it} - m_{it}$, and $y^*_{it} - y_{it}$ prior to estimating (3). Thus, as with the dynamic SUR estimator, our ARDL estimator does not estimate the cointegrating coefficients for each country in isolation. Again, we consider this estimator in order to establish a basis for another test of the homogeneity restrictions inherent in panel estimators. ARDL parameter estimates for β_{i1} and β_{i2} are reported in columns (6) and (7) of Table 2.[9] The ARDL estimator generally generates coefficient estimates that are more in line with the monetary model than the previous five estimators, as the β_{i1} and β_{i2} estimates both have the correct sign for 11 of the 18 countries. However, the standard errors are typically fairly large, and only for Japan are both coefficient estimates correctly signed and significantly different from zero. And even for Japan, the β_{i1} estimate of 2.19 is "too large" in that we reject the null hypothesis that β_{i1} is unity, as required by the monetary model.

In summary, none of the six estimators that we consider in Tables 1 and 2 generate more than slight support for the monetary model. These results reinforce those from other studies that test the monetary model on a country-by-country basis using post-Bretton Woods data. Our results add value by clearly demonstrating the extent of the lack of support for the monetary model on a country-by-country basis during the modern float.[10]

[9] We select p, q_1, and q_2 in eq. (3) using the AIC.

[10] At the suggestion of a referee, we tested the stability of the cointegrating vectors using the Hansen and Johansen (1993) recursive likelihood-ratio statistics. There is extensive evidence of structural instability in the cointegrating vectors according to the recursive likelihood-ratio statistics. We also tested the stability of the cointegrating vectors using the *SupF* and *AveF* statistics and the distributions in Hansen (1992). There is evidence of instability in the cointegrating vector for a number of countries according to the *SupF* and *AveF* statistics. Parameter instability is thus one potential reason for the failure of the cointegrating parameters to conform to the theoretical priors. All unreported results are available upon request from the authors.

2.4 Country-by-Country Cointegration Test Results

Testing for the existence of a cointegrating relationship on a country-by-country basis is anticlimactic given that, even when we assume the existence of cointegrating relationship among nominal exchange rates and relative money supplies and income levels, we do not obtain cointegrating coefficient estimates consistent with the monetary model. Nevertheless, in order to allow a full comparison of the country-by-country test results with our panel test results, we use the two most popular extant procedures—those of Engle and Granger (1987) and Johansen (1991)—to test for cointegration on a country-by-country basis.

The well-known augmented Engle and Granger (1987, AEG) procedure examines the stationarity of the OLS residuals from the estimated cointegrating relationship (2) for country i. In particular, an augmented Dickey and Fuller (1979) regression is applied to the OLS residuals $\hat{u}_{it}$ for country i:

$$\Delta \hat{u}_{it} = \rho_i \hat{u}_{i,t-1} + \sum_{j=1}^{k} \Delta \hat{u}_{i,t-j} + \epsilon_{it}. \tag{4}$$

A constant term is not included in (4), as it is already included in (2). Under the null hypothesis of no cointegration, the residuals from (2) contain a unit root so that $\rho_i = 0$. Under the alternative hypothesis, the residuals are stationary so that $\rho_i < 0$. We can test the null hypothesis by inspecting the t-statistic corresponding to ρ_i in (4). Column (2) of Table 3 reports AEG test results on a country-by-country basis.[11] The AEG null hypothesis of no cointegration cannot be rejected for any of the 18 countries.

We also report results for the widely used Johansen (1991) trace test in columns (3) and (4) of Table 3. Like the JOH-ML estimates, the trace statistic is derived from maximum-likelihood estimation of a vector-error-correction model,[12] and like the AEG test, no cointegration serves as the null hypothesis for the trace test. From column (3) of Table 3, we see that the trace test rejects the null of no cointegration in favor of the alternative of cointegration for 11 (13) countries at the 5 (10) percent significance level. However, the results using the conventional trace statistics should be viewed with some suspicion, as Cheung and Lai (1993) show that the conventional trace statistic can be seriously oversized. When we use the Cheung and Lai (1993) degrees-of-freedom correction, the trace test is only significant at the 5 (10) percent level for

[11] We select the lag order k in (4) using the Campbell and Perron (1991) top-down procedure.

[12] As with the JOH-ML estimates, we select the lag order for the vector-error-correction model using top-down testing and the Sims (1980) modified likelihood-ratio statistic.

four (six) countries.[13] Even for the six countries where the degress-of-freedom corrected trace statistic is significant at the 10 percent level, this is not sufficient evidence for the monetary model, since (as we emphasized above) the JOH-ML coefficient estimates reported in Table 2 are not in accord with the monetary model for these six countries. Note that our results for Canada are similar to those reported in Cushman (2000). Using the Johansen (1991) procedure, Cushman (2000) finds evidence of a cointegrating relationship among nominal U.S. dollar-Canadian dollar exchange rates and a set of monetary fundamentals over the post-Bretton Woods float, but the estimated cointegrating coefficients are not at all in line with the monetary model. We show that a similar result holds for Denmark, Germany, Japan, the Netherlands, and Sweden.

As we noted in the introduction, it is difficult to overstate how poorly the monetary model fails on a country-by-country basis during the post-Bretton Woods float. Given the paucity of evidence for the monetary model on a country-by-country basis, we next turn to panel cointegration tests.

3 Panel Results

Similar to the Engle and Granger (1987) framework, panel cointegration tests typically involve testing the stationarity of the residuals from a levels regression. However, panel techniques may be better able to detect cointegrating relationships since a pooled levels regression exploits cross-sectional information in the data when estimating cointegrating coefficients. Furthermore, the number of observations available when testing the stationarity of the residual series from the levels regression is greatly increased in a panel framework, and this can substantially increase the power of cointegration tests. This is especially important in the present context, given the notoriously low power of country-by-country cointegration tests against persistent but stationary alternatives for samples as short as the approximately 25 years making up the modern float. We present a number of panel estimates of cointegrating coefficients in Section 3.1, and we report results of residual-based panel tests for the existence of a cointegrating relationship among nominal exchange rates, relative money supplies, and relative income levels in Section

[13] We make inferences similar to those for the degrees-of-freedom corrected trace statistic using a bootstrap procedure like the one in Cushman (2000).

3.2.[14]

3.1 Panel Cointegrating Coefficient Estimates

We first consider the LSDV estimator (also known as the pooled fixed-effect estimator), which entails OLS estimation of (2) under the set of homogeneity restrictions, $\beta_{i1} = \beta_1$ and $\beta_{i2} = \beta_2$ for $i = 1, \ldots, N$:

$$e_{it} = \beta_{i0} + \beta_1(m_{it}^* - m_{it}) + \beta_2(y_{it}^* - y_{it}) + u_{it}. \tag{5}$$

LSDV estimates for our full panel of 18 countries and four subpanels are reported in columns (2) and (3) in the top half of Table 4. All five of the β_1 LSDV estimates and four of five of the β_2 LSDV estimates have the correct sign. Unfortunately, we cannot use the conventional LSDV standard errors for valid inference, as Kao and Chiang (2000) show that the LSDV estimates suffer from asymptotic bias (similar to OLS estimation of a cointegrating relationship for a single cross-sectional unit).[15] Fortunately, Kao and Chiang (2000) also develop a bias-corrected LSDV estimator that can be used for inference, and bias-corrected LSDV estimates are reported in columns (4) and (5) in the top half of Table 4.[16] We see that the bias-adjusted LSDV point estimates are very similar to the unadjusted LSDV point estimates. However, the standard errors for the bias-adjusted LSDV estimates are generally larger than those for the unadjusted LSDV estimates, reflecting the bias in the unadjusted LSDV standard errors. All of the β_1 bias-adjusted LSDV estimates have the correct sign, and, with the exception of the Group of 6 subpanel, all of these estimates are significant at the 5 percent level. However, all of the β_1 estimates are significantly less than unity, the β_1 value required by the monetary model. With the exception of the EMS subpanel, all of the β_2 bias-adjusted LSDV estimates have the correct sign, and the β_2 estimates are significant for the Group of 6 and Group of 10 subpanels and the full panel.

Both LSDV estimators in the top half of Table 4 implicitly assume that the error terms in (5) are independent across the cross-sectional units. One way to handle cross-sectional correlation in the error terms is through SUR estimation of the cointegrating relationship (5). In columns

[14] See Banerjee (1999) and Baltagi and Kao (2000) for surveys of panel unit root and cointegration tests.

[15] A type of super-consistency also obtains for the LSDV estimator. Kao and Chiang (2000) use the sequential limit theory of Phillips and Moon (1999) in which $T \to \infty$ followed by $N \to \infty$.

[16] We use the GAUSS program NPT 1.2 (authored by Min-Hsien Chiang and Chihwa Kao) available from Chihwa Kao's home page (http://www.maxwell.syr.edu/maxpages/faculty/cdkao/working/w.html) to generate the bias-corrected LSDV estimates.

(6) and (7) of the top half of Table 4, we report Mark, Ogaki, and Sul (2000) restricted dynamic SUR estimates of β_1 and β_2. These are calculated the same way as the unrestricted dynamic SUR estimates reported in Table 2, but with the restriction that the slope coefficients are homogeneous across countries imposed. The restricted dynamic SUR estimates perform reasonably well in terms of the monetary model, as the β_1 and β_2 estimates have the correct signs for the full panel and each of the four subpanels. In addition, all five of the β_1 estimates are significant, and four of the five β_2 estimates are significant. However, all of the β_1 estimates are significantly below unity.

Overall, the panel estimates in the top half of Table 4 provide more support for the monetary model than the country-by-country estimates reported in Tables 1 and 2 in that the β_1 and β_2 panel estimates almost always have the correct sign and are often significant. However, the β_1 panel estimates in the top half of Table 4 are generally significantly less than the value of one required by the monetary model. Groen (2000, Table 3) reports LSDV estimates of (5) for a panel of 14 industrialized countries using data spanning 1973:1–1994:4. The β_1 and β_2 estimates all have the correct signs for his full panel and three subpanels. The β_1 estimates are all below one, ranging from 0.664 for the full panel to 0.832 for a Group of 10 subpanel. However, Groen (2000) does not report standard errors for the point estimates, so one cannot assess the significance of the point estimates and whether the β_1 estimates are significantly different from unity. In summary, it appears that panel estimates of the cointegrating coefficients β_1 and β_2 that do not account for correlation across error terms or that account for correlation via SUR produce β_2 estimates that are significantly negative (as required by theory) and β_1 estimates that are significantly positive but also significantly less than the value of unity required by theory.

In columns (2)–(5) in the bottom half of Table 4, we report LSDV and bias-adjusted LSDV coefficient estimates that account for cross-sectional error correlation by allowing for a common time effect. When allowance is made for a common time effect, (5) becomes:

$$e_{it} = \beta_{i0} + \theta_t + \beta_1(m_{it}^* - m_{it}) + \beta_2(y_{it}^* - y_{it}) + v_{it}. \tag{6}$$

This is a popular method of accounting for cross-section correlation, as it is easily implemented by extracting each period's cross-sectional mean from each variable prior to estimating the cointegrating relationship. One can see from the bottom half of Table 4 that the unadjusted and bias-adjusted LSDV point estimates are similar. We concentrate on the bias-adjusted estimates in what follows, as we can use the bias-adjusted standard errors for valid inference. All five of the β_1 estimates have the correct sign and are significant. In addition, none of the β_1 estimates are significantly different from unity. All five of the β_2 estimates have the correct sign, and all are significant. Three of the β_2 estimates are not significantly different from -1, in support of a money demand income elasticity of unity. Mark and Sul (2001) assume cointegrating coefficients of $\beta_1 = 1$ and $\beta_2 = -1$ in their panel cointegration test, and our estimation results provide some support for this assumption. Overall, LSDV estimates that allow for a common time effect perform remarkably well in terms of the monetary model and stand in stark contrast to the country-by-country estimates.

Finally, β_1 and β_2 estimates generated using the Pesaran, Shin, and Smith (1999) pooled mean group estimator with a common time effect are reported in columns (6) and (7) in the bottom half of Table 4. Pesaran, Shin, and Smith (1999) assume an ARDL specification for each country with the long-run (cointegrating) coefficients restricted to be homogeneous across countries, and the pooled mean group estimator is generated through maximum-likelihood estimation of the restricted ARDL system.[17] From columns (6) and (7) in the bottom half of Table 4, we see that all of the β_1 and β_2 estimates have the correct signs in terms of the monetary model. Three of the five β_1 estimates are significant, and the β_1 estimates for the Group of 10 subpanel and the full panel are not significantly different from unity. Three of the five β_2 estimates are also significant, although the significant β_2 estimate for the Group of 10 subpanel appears somewhat large (in absolute value) when interpreted as a money demand income elasticity. The pooled mean group estimates for the full panel are very much in line with the monetary model: the β_1 estimate of 0.90 is statistically significant but not significantly different from unity, and the β_2 estimate of -0.73 is statistically significant and plausible in magnitude.

Why are the panel estimates of the cointegrating coefficients so much more plausible in terms of the monetary model than the country-by-country estimates? Of course, there is a substantial increase in the sample size when we pool the data, and (ceteris paribus) this increases the

[17]The short-run coefficients are free to vary across countries. We use the GAUSS program (authored by Yongcheol Shin) available from M. Hashem Pesaran's home page (http://www.econ.cam.ac.uk./faculty/pesaran) to generate the pooled mean group estimates.

precision of the estimates. More variation in the regressors also increases the precision of the estimates. In Table 5, we compare the country-by-country and panel variation in the regressand and regressors when estimating the cointegrating coefficients. Column (4) reports the average standard deviation for the relative money supply on a country-by-country basis, while column (5) reports the standard deviation for the panel. We see that the variation in the relative money supply regressor in much greater for the panels than for each country on average, with the panel standard deviations all at least an order of magnitude greater than the country-by-country averages. The variation in the relative real output level regressor is also greater for the panels, but the increase is not as marked as for the relative money supplies. The increased number of observations and increased variability in the regressors apparently enables the panel estimates to converge on the hypothesized coefficient values, providing much more support for the monetary model. Note, however, that in order to take advantage of the increases in sample size and regressor variation that result from pooling, the homogeneity assumption that $\beta_{i1} = \beta_1$ and $\beta_{i2} = \beta_2$ in (2) must be valid. We investigate this homogeneity assumption in detail in Section 4 below.

3.2 Panel Residual-Based Cointegration Test Results

While a number of the panel cointegrating coefficient estimates reported in Table 4 are in line with the monetary model, we need to test for the stationarity of the panel regression residuals, as this is also a necessary condition for the monetary model. We consider five panel residual-based cointegration tests. As Groen (2000) and Mark and Sul (2001) consider only a single residual-based test each, we can examine the robustness of their cointegration inferences. Analogous to the Engle and Granger (1987) and Phillips and Ouliaris (1990) two-step cointegration tests, all five of our tests examine the stationarity of the residuals from an estimated cointegrating relationship.

Our first three tests are based on the residuals from the LSDV regression model with a common time effect (6). These three tests thus account for cross-sectional dependency in the disturbances by allowing for a common time effect in the first step of the test (where the cointegrating relationship is estimated). The first test is due to Pedroni (1995) and is essentially a pooled version of the Phillips and Ouliaris (1990) cointegration test. Let $\hat{v}_{it}$ ($i = 1, \ldots, N$, $t = 1, \ldots, T$) be the estimated residuals from the first-step regression (6) and consider the following second-step regression model:[18]

$$\Delta \hat{v}_{it} = \rho \hat{v}_{i,t-1} + \eta_{it}. \tag{7}$$

Under the null of no cointegration, the residuals from (6) contain a unit root so that $\rho = 0$, while under the alternative hypothesis of cointegration the residuals are stationary so that $\rho < 0$. Pedroni (1995) first defines $\hat{\rho}$ as:

$$\hat{\rho} = \left(\sum_{i=1}^{N} \sum_{t=1}^{T} \hat{v}_{i,t-1}^2 \right)^{-1} \sum_{i=1}^{N} \sum_{t=1}^{T} \left(\hat{v}_{i,t-1} \Delta \hat{v}_{it} - \hat{\lambda}_i \right), \tag{8}$$

where $\hat{\lambda}_i = T^{-1} \sum_{s=1}^{k} w_{sk} \sum_{t=1}^{T} \hat{\omega}_{it} \hat{\omega}_{i,t-s}$ for some lag window w_{sk} and where $\hat{\omega}_{it}$ are the estimated residuals from the autoregression $\hat{v}_{it} = \delta \hat{v}_{i,t-1} + \omega_{it}$. What we label the PC statistic, PC $= \sqrt{NT(T-1)}\hat{\rho}$, can then be used to test the null hypothesis that $\rho = 0$ against the alternative that $\rho < 0$ in (7). It is clear that the Pedroni (1995) PC statistic accounts for serial correlation using a semi-parametric adjustment. In our applications, we use the Bartlett kernel and a lag truncation of eight for w_{sk}.[19] We consider two additional tests of the null hypothesis that $\rho = 0$ against the alternative that $\rho < 0$ in (7) by employing the Kao (1999) DF_ρ^* and DF_t^* statistics. The details of the construction of the DF_ρ^* and DF_t^* statistics are given in Kao (1999).[20]

Our final two panel cointegration tests are the Groen (2000) test and a Taylor and Sarno (1998) modification of the Groen (2000) test. The Groen (2000) test uses the estimated residuals from the LSDV regression without common time effects, (5), so that cross-sectional correlation in the disturbance terms is not accounted for in the test's first step. Letting $\hat{u}_{it}$ be the estimated residuals from (2), Groen (2000) considers the following panel augmented Dickey and Fuller (1979) regression for $i = 1, \ldots, N$:

$$\Delta \hat{u}_{it} = \rho \hat{u}_{i,t-1} + \sum_{j=1}^{p} \phi_{ij} \Delta \hat{u}_{i,t-j} + \nu_{it}, \tag{9}$$

[18] We do not include a constant in (7), as a constant is already included in the first-step regression (6).

[19] The results are not sensitive to the kernel function.

[20] Kao (1999) derives three other statistics (DF_ρ, DF_t, and ADF) that can also be used to test the null hypothesis that $\rho = 0$ in (7). We use the DF_ρ^* and DF_t^* statistics, as these have the best overall performance in the Monte Carlo simulations conducted by Kao (1999). The Kao (1999) DF_ρ^* and DF_t^* statistics are similar to the PC statistic in that they involve semi-parametric adjustments. As with the PC statistic, we use the Bartlett kernel and a lag truncation of eight for the DF_ρ^* and DF_t^* statistics.

so that a parametric procedure is used to account for serial correlation. Equation (9) is estimated using feasible GLS (SUR), so that cross-sectional correlation is accounted for in the test's second step. Under the null hypothesis of no cointegration, the residuals from (5) are nonstationary so that $\rho = 0$ in (9). Using the t-statistic on ρ in (9), the null of no cointegration is tested against the alternative of cointegration ($\rho < 0$). The lag order p in (9) is selected using top-down testing.[21] We label the t-statistic on ρ in (9) the SUR pooled ADF (SURPADF) statistic. We also consider a modification of the Groen (2000) test analogous to the Taylor and Sarno (1998) MADF test. In order to understand the modification, note that all of the panel cointegration tests we have considered so far employ pooled estimates of ρ so that they impose the restriction that $\rho_i = \rho$ in (7) or (9). This means that the autoregressive parameter is restricted to be identical across countries under both the null and alternative hypotheses. Our modified Groen (2000) test allows the autoregressive parameter to vary across countries under the alternative hypothesis so that (9) becomes:

$$\Delta \hat{u}_{it} = \rho_i \hat{u}_{i,t-1} + \sum_{j=1}^{p} \phi_{ij} \Delta \hat{u}_{i,t-j} + \nu_{it}. \tag{10}$$

Following Taylor and Sarno (1998), we then test the null hypothesis that $\rho_i = 0$ for all i using a Wald test. We again select the lag order p using top-down testing, and we label the Wald statistic used to test the null that $\rho_i = 0$ for all i the MADF statistic (following Taylor and Sarno 1998).[22]

In order to increase the likelihood of making accurate inferences, we base all of our panel cointegration test inferences on bootstrapped critical values. We use a nonparametric version of the parametric bootstrap procedure in Groen (2000) to generate critical values for our five tests. In contrast to the Groen (2000) parametric bootstrap, our nonparametric procedure does not depend on the assumption that the residuals are normally distributed, and it allows for a greater degree of contemporaneous correlation across nominal exchange rates, relative money

[21]Following Groen (2000), the top-down test proceeds as follows. Starting with $p_{max} = 5$, we estimate (9) via feasible SUR. We then jointly test the significance of $\phi_{15}, \ldots, \phi_{N5}$ using a likelihood-ratio test. If we reject this null, we set $p = 5$ in (9). If the null is rejected, we estimate (9) with $p = 4$ and test the null hypothesis that $\phi_{14}, \ldots, \phi_{N4} = 0$. We continue in this manner until we reject the null hypothesis. If none of the null hypotheses for $p = 1, \ldots, 5$ is rejected, we set $p = 0$.

[22]Note that rejection with the MADF test does not imply that cointegration holds for all panel members, as it admits the possibility that cointegration holds for some, but not necessarily all, panel members under the alternative hypothesis. The MADF test thus is limited in how informative it is; see Breuer, McNown, and Wallace (2001).

supplies, and relative income levels. Like Groen (2000), our bootstrap is based on the following data generating process:

$$e_{it} = \beta_{i0} + \beta_1(m^*_{it} - m_{it}) + \beta_2(y^*_{it} - y_{it}) + \xi_{it}, \tag{11}$$

$$\Delta\xi_{it} = \textstyle\sum_{i=1}^{p} \tau_{i1j}\Delta\xi_{i,t-j} + \pi^{\xi}_{it}, \tag{12}$$

$$\Delta(m^*_{it} - m_{it}) = \tau_{i20} + \tau_{i21}\Delta(m^*_{i,t-1} - m_{i,t-1}) + \pi^{m}_{it}, \tag{13}$$

$$\Delta(y^*_{it} - y_{it}) = \tau_{i30} + \tau_{i31}\Delta(y^*_{i,t-1} - y_{i,t-1}) + \pi^{y}_{it}. \tag{14}$$

Note that the ξ_{it} term in (11) is nonstationary according to the specification in (12), so that e_{it}, $m^*_{it} - m_{it}$, and $y^*_{it} - y_{it}$ are not cointegrated. This is in accord with the null hypothesis for each of our panel cointegration tests in Section 3.2.

As in Groen (2000), we first estimate (11) via LSDV using the original data. This gives us the parameter estimates $\hat{\beta}_{i0}$, $\hat{\beta}_1$, and $\hat{\beta}_2$, as well as a set of residual series $\hat{\xi}_{it}$. We then use the $\hat{\xi}_{it}$ series to estimate (12) for each country, yielding a set of parameter estimates $\hat{\tau}_{i11}$ through $\hat{\tau}_{i1p}$ and another set of residual series $\hat{\pi}^{\xi}_{it}$. We also estimate (13) and (14) for each country, in order to obtain the parameter estimates $\hat{\tau}_{i20}$, $\hat{\tau}_{i21}$, $\hat{\tau}_{i30}$, and $\hat{\tau}_{i31}$ and the residual series $\hat{\pi}^{m}_{it}$ and $\hat{\pi}^{y}_{it}$ for each country.

For each time period t ($t = 1, \ldots, T$), the residual series $\hat{\pi}^{\xi}_{it}$, $\hat{\pi}^{m}_{it}$, and $\hat{\pi}^{y}_{it}$ ($i = 1, \ldots, N$) can be organized into a $3 \cdot N$-vector of residuals. In order to generate pseudo-disturbances for the bootstrap procedure, we randomly draw (with replacement) one complete vector at a time from the T residual vectors. Note that this procedure does not rely on the normality of the residuals, and it preserves the correlation structure of the original data across all variables and all countries.[23] Armed with the estimates of the parameters in (11)–(14) and the vectors of pseudo-disturbances ($\tilde{\pi}^{\xi}_{it}$, $\tilde{\pi}^{m}_{it}$, and $\tilde{\pi}^{y}_{it}$ for $i = 1, \ldots, N$ and $t = 1, \ldots, T$), we can generate a pseudo-sample of observations for e_{it}, $m^*_{it} - m_{it}$, and $y^*_{it} - y_{it}$ using the data generating process (11)–(14). The initial values of $\xi_{i,t-j}$, $\Delta(m^*_{i,t-1} - m_{i,t-1})$, and $\Delta(m^*_{i,t-1} - m_{i,t-1})$ are set to zero. We generate a pseudo-sample of $T + 100$ observations, and we discard the first 100 observations in order to randomize the initial values. We repeat this process 500 times, calculating and storing the five statistics of Table 5 for each replication. The stored statistics are then sorted.

[23]The Groen (2000) parametric bootstrap procedure assumes that the residuals are normally distributed, and it only accounts for correlation across the $\hat{\pi}^{\xi}_{it}$ residuals.

For the PC, DF_ρ^*, DF_t^*, and SURPADF statistics, the fifth, 25th, and 50th sorted statistics serve as the 1, 5, and 10 percent critical values; for the MADF statistic, the 495th, 475th, and 450th statistics serve as the 1, 5, and 10 percent critical values.

Table 6 reports our panel cointegration test results. We see from the table that there is considerable evidence of cointegration. All five statistics allow us to reject the null hypothesis of no cointegration in favor of the alternative of cointegration at the 10 percent level or better for the EC and Group of 6 subpanels and the full panel. Most of these rejections occur at the 5 percent level or lower. Four of the five tests reject no cointegration for the Group of 10 subpanel, while there are no rejections for the EMS subpanel. Overall, there is substantial evidence of cointegration among nominal exchange rates, relative money supplies, and relative income levels in Table 6, in contrast to the country-by-country cointegration test results reported in Table 3.[24]

Recall that the PC, DF_ρ^*, and DF_t^* statistics are based on the residuals from (6), the LSDV regression with a common time effect. Given that the LSDV estimates of β_1 and β_2 from (6) reported in Table 4 are very much in line with the values predicted by the monetary model and given that the PC, DF_ρ^*, and DF_t^* statistics typically reject the null hypothesis of no cointegration in favor of cointegration, we have strong support for the monetary model when we allow for a common time effect. Mark and Sul (2001) assume cointegrating coefficients of $\beta_1 = 1$ and $\beta_2 = -1$ and find evidence of cointegration using a test that allows for a common time effect. Our results generally support those in Mark and Sul (2001), as most of our LSDV estimates of β_1 and β_2 from (6) reported in Table 4 are reasonably consistent with $\beta_1 = 1$ and $\beta_2 = -1$, and our three panel tests based on the residuals from (6) indicate cointegration.

4 To Pool or Not to Pool?

As a comparison of the results reported in Sections 2 and 3 makes clear, there is much more support for the monetary model if we pool the data and use panel procedures. Since we find much more support for the monetary model by pooling, it is natural to inquire whether the

[24]At the suggestion of a referee, we also tested for cointegration using heterogeneous panel tests based on the residuals from (2), so that the homogeneity restrictions in (5) are not imposed (see Pedroni 1999). Similar to the results in Table 6, there is considerable evidence of cointegration for the various panels using heterogeneous panel tests. However, this alone should not be construed as support for the monetary model, as the heterogeneous cointegrating parameter estimates in Tables 1 and 2 are typically not in line with the monetary model.

homogeneity restrictions inherent in pooling are consistent with the data. Casual inspection of the country-by-country cointegrating coefficient estimates in Tables 1 and 2 suggests that the coefficients are not homogeneous, as they differ widely across countries for most estimators. We consider two formal tests of the homogeneity restrictions, $\beta_{i1} = \beta_1$ and $\beta_{i2} = \beta_2$ for $i = 1, \ldots, N$. The first is a Wald test due to Mark, Ogaki, and Sul (2000), and it can be used to test the null hypothesis that the homogeneity restrictions hold across the unrestricted dynamic SUR estimates in columns (4) and (5) of Table 2. Our second test is a likelihood-ratio test of the null hypothesis that the homogeneity restrictions hold across the ARDL estimates in columns (6) and (7) of Table 2. This latter test is from Pesaran, Shin, and Smith (1999).

We report the results for the two tests of the homogeneity restrictions in Table 7. The Wald statistic rejects the homogeneity restrictions at the 1 percent level for the full panel and three of the four subpanels, and the homogeneity restrictions are rejected at the 5 percent level for the other subpanel. The likelihood-ratio statistic rejects the null of homogeneity at the 1 percent level for the full panel and the EC subpanel, and it rejects at the 5 percent level for the Group of 6 subpanel. The homogeneity restrictions are not rejected for the EMS and Group of 10 subpanels at conventional significance levels, but the p-values are still quite close to 0.10 so that support for the restrictions is not strong. Overall, the results in Table 6 indicate that the data provide little support for the homogeneity assumptions inherent in panel estimators of the monetary model, including the homogeneity assumptions in Groen (2000) and Mark and Sul (2001).

The lack of statistical support for the homogeneity assumptions leaves us in a quandary. When we test the monetary model on a country-by-country basis, there is essentially no support for the monetary model. Conversely, we find considerable support for the monetary model when we pool the data. Ceteris paribus, we would like to take advantage of the additional increase in sample size and regressor variation that results from pooling. However, formal statistical tests typically do not support pooling. If the coefficients are truly heterogeneous across countries, we run the risk of obtaining spurious evidence of cointegration by applying panel tests. Observe that if a given researcher has the strong prior belief that homogeneity restrictions should only be imposed if they pass formal statistical tests, then the researcher will not find support for the monetary model in post-Bretton Woods data.[25]

[25] One can find some evidence in support of the monetary model for a small panel of countries that does not

Interestingly, Pesaran, Shin, and Smith (1999) obtain similar results in their panel study of consumption and energy demand functions. They find that some individual countries have long-run consumption or energy demand income elasticities that are at odds with theory, while their pooled mean group estimates of the long-run coefficients are more sensible in terms of theory. However, likelihood-ratio tests reject the homogeneity restrictions inherent in the pooled mean group estimates. Pesaran, Shin, and Smith (1999) offer a possible explanation for the rejection of the homogeneity restrictions: the coefficient estimates of individual members of the panel may be biased because of sample-specific omitted variables or measurement errors that are correlated with the regressors. If the coefficients are, in fact, the same across the individual cross-sectional units, and the correlations responsible for the bias are not systematic (so that they average out to zero), then pooling is appropriate. In this case, the pooled estimates are informative and have economic content despite the rejection of the homogeneity restrictions. Unfortunately, as they note, "there is no obvious way of determining from the data that is the case." We observe that if a researcher has the strong prior belief that pooling successfully eliminates the biases in country-by-country estimates of the monetary model, then the research will find substantial support for the monetary model in post-Bretton Woods data.

Baltagi, Griffin, and Xiong (2000) is another interesting parallel in a microeconomic context. Baltagi, Griffin, and Xiong (2000) investigate cigarette demand for a panel of U.S. states. They compare a large number of pooled and heterogeneous estimators, asking the question, "To pool or not to pool?" The pooled estimates are in line with theory, while some of the heterogeneous estimates are implausible (for example, positive own-price elasticities). A Wald test rejects the null hypothesis that the slope coefficients are the same across the cross-sectional units. Despite the rejection, Baltagi, Griffin, and Xiong (2000) recommend pooling, as the pooled estimates generate superior out-of-sample forecasts (in addition to being more plausible in terms of theory). Baltagi, Griffin, and Xiong (2000) argue that the superior forecast performance of the pooled estimates make them more reliable.

We perform a similar out-of-sample forecasting exercise using our panel and country-by-

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entail a rejection of the homogeneity restrictions. Looking back to Table 1, we see that the OLS estimates of the cointegrating coefficients have the correct sign for Australia, France, Italy, and Spain. If we combine these four countries into a panel, the Mark, Ogaki, and Sul (2000) Wald test does not reject the homogeneity restrictions. The LSDV estimates of β_1 and β_2 for this panel are 0.59 and -0.48, which are somewhat consistent with the monetary model, and the SURPADF and MADF statistics reject the null hypothesis of no cointegration at the 10 percent level. We thank a referee for suggesting this panel.

country estimates. In order to compare forecast performance, we compare out-of-sample forecasts generated through a predictive regression, as in Mark (1995). We first form nominal exchange rate forecasts based on the country-by-country estimates using the following predictive regression:

$$\Delta e_{i,t+k} = \kappa_{i0} + \kappa_{i1} z_{it} + \phi_{it}, \tag{15}$$

where z_{it} is the deviation of the nominal exchange rate from its value predicted by the levels regression (2). That is,

$$z_{it} = e_{it} - [\hat{\beta}_{i1}(m_{it}^* - m_{it}) + \hat{\beta}_{i2}(y_{it}^* - y_{it})], \tag{16}$$

where $\hat{\beta}_{i1}$, $\hat{\beta}_{i2}$ are individual country estimates of β_{i1} and β_{i2}. Using the first 50 sample observations, we estimate (15) by OLS in order to generate an out-of-sample k-step-ahead forecast for Δe_{50+k}. We then update the sample by one observation, again estimate (15), and generate a forecast for Δe_{51+k}. We continue this process through the end of the available sample for each country. For each forecast for each country, we compute the forecast error, and we combine all the forecast errors in order to compute the root mean-squared prediction error (RMSE) for the forecasts corresponding to the country-by-country or heterogeneous estimates. We also generate out-of-sample forecasts corresponding to the pooled or homogeneous estimates using (16) but with $\hat{\beta}_{i1} = 1$ and $\hat{\beta}_{i2} = -1$ for $i = 1, \ldots, 18$. The LSDV β_1 and β_2 pooled estimates that allow for a common time effect are quite close to—and not significantly different from—the values of 1 and -1, so we take this to be an accurate characterization of the homogeneous estimates.[26]

A comparison of the heterogeneous and homogeneous forecasts is presented in Table 8, which reports the ratio of the RMSE corresponding to the homogeneous estimates (where $\beta_1 = 1$ and $\beta_2 = -1$) to the RMSE corresponding to the heterogeneous estimates. We consider each of the heterogeneous estimates reported in Tables 1 and 2.[27] The 1-step-ahead and 4-step-ahead RMSEs for the homogeneous and heterogeneous estimates are very similar. At the 8-step-ahead horizon, the homogeneous estimates generate better forecasts in comparison to five of the six heterogeneous estimates. At the 16-step-ahead horizon, the homogeneous estimates have an

[26] Using β_1 and β_2 values of 1 and -1 also facilitates comparison to Mark and Sul (2001).

[27] We consider forecasts at horizons of 1, 4, 8, 12, and 16 quarters. Given that we use the first 50 observations to generate the first forecast, we are able to compute 43, 40, 36, 32, and 28 forecasts for each country at the 1-, 4-, 8-, 12-, and 16-step-ahead horizons.

RMSE that is smaller than the RMSE for each of the heterogeneous estimates. In most cases, the RMSE is reduced by nearly 20 percent when the homogeneous estimates are used at the 16-step-ahead horizon.[28] If we adopt the stance of Baltagi, Griffin, and Xiong (2000), we may favor the pooled estimates of the monetary model over the country-by-country estimates, as the pooled estimates are much more plausible in economic terms, and they generate superior forecasts at longer horizons.

Coakley, Fuertes, and Smith (2001) raise another issue relating to pooling. They analyze the small-sample properties of pooled estimators under a variety of data generating processes through Monte Carlo experiments. They find that pooled estimators can lead to distorted inferences when the true coefficients are heterogeneous. We undertake a Monte Carlo experiment similar in spirit to those in Coakley, Fuertes, and Smith (2001). Recall the nonparametric bootstrap procedure that we used to generate critical values for the panel cointegration tests. We use the same data generating process, with the exception that the country-by-country OLS estimates from Table 2 are used instead of the LSDV (homogeneous) estimates in (11). We thus have a data generating process in which the underlying coefficients are heterogeneous (and not in line with the monetary model) and the disturbances are $I(1)$ in (11). We generate 500 pseudo-samples of panel data and calculate LSDV estimates that allow for a common time effect for each replication in order to create empirical distributions for the LSDV estimates of β_1 and β_2 that (wrongly in this case) assume homogeneous coefficients. Table 9 reports summary statistics for the empirical distributions. Interestingly, the summary statistics in part A of Table 9 suggest that it would not be unlikely for us to obtain the LSDV β_1 estimates reported in column (2) of the bottom half of Table 4 (which are reasonably in line with the monetary model) even if the true data generating process is not in accord with the monetary model. As similar result holds for the LSDV β_2 estimates, although it appears to be somewhat less likely that we would obtain the LSDV β_2 estimates in column (3) of the bottom half of Table 4 if the true data generating process does not accord with the monetary model. Overall, simulations along the lines of Coakley, Fuertes, and Smith (2001) suggest that pooling the data when the true data generating process is characterized by heterogeneity can produce spurious evidence in support of the monetary model.

[28] We stress that this is a simple mechanical comparison of forecasts, as in Baltagi, Griffin, and Xiong (2000). Berben and van Dijk (1998), Kilian (1999), and Berkowitz and Giorgianni (2001) raise a number of key issues relating to the use of predictive regressions to assess the ability of fundamentals to predict exchange rates.

In summary, there are good reasons to favor the panel estimates over the country-by-country estimates of the monetary model, but there are also good reasons to be suspect of the panel estimates.

5 Conclusion

In this paper, we take a closer econometric look at panel tests of the monetary model. We show that, in stark contrast to country-by-country tests, there is substantial support for the monetary model using panel tests. In particular, a number of panel estimates of the cointegrating coefficients are very much in line with the monetary model, and panel cointegration tests largely indicate that U.S. dollar exchange rates, relative money supplies, and relative income levels are cointegrated. However, an additional finding of ours makes it very difficult to assess the extent of support for the monetary model in post-Bretton Woods data: while we show that it is necessary to make homogeneity assumptions concerning cointegrating coefficients in order to find support for the monetary model, such assumptions are not supported by the data in that formal tests of the homogeneity restrictions are rejected. Interestingly, similar results appear in Pesaran, Shin, and Smith (1999) and Baltagi, Griffin, and Xiong (2000) in different contexts. Like Pesaran, Shin, and Smith (1999) and Baltagi, Griffin, and Xiong (2000), we believe that panel estimates should not necessarily be dismissed on the basis of tests of homogeneity restrictions alone. Pesaran, Shin, and Smith (1999) argue that panel estimation procedures may eliminate certain biases that plague country-by-country estimates. A comparison of forecasts in the spirit of Baltagi, Griffin, and Xiong (2000) also indicates that panel estimates generate out-of-sample nominal exchange rate forecasts that are superior to those generated by country-by-country estimates, suggesting that panel estimates of the monetary model are more reliable. However, a Monte Carlo experiment shows that it is not improbable to find evidence in support of the monetary model by relying on panel estimates, even though the true data generating process is characterized by a heterogeneous structure that is not consistent with the monetary model. Given that homogeneity restrictions are rejected by formal tests, this latter finding suggests that the support for the monetary provided by panel procedures model is spurious. In the end, it is quite difficult to say whether post-Bretton Woods data conforms to the monetary model, and a given researcher's final assessment will largely depend on the researcher's prior beliefs concerning the following: the role of formal tests of homogeneity restrictions; the possibility that biases in country-by-country estimates are "averaged out" in panel estimates; and the role of out-of-sample forecasts in determining the reliability of panel estimates.

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Table 1: Country-by-country OLS-based estimates of the cointegrating relationship, $e_{it} = \beta_{i0} + \beta_{i1}(m^*_{it} - m_{it}) + \beta_{i2}(y^*_{it} - y_{it})$

(1)	(2)	(3)	(4)	(5)	(6)	(7)
	OLS estimates		FM-OLS estimates		DOLS esitmates	
Country	β_{i1}	β_{i2}	β_{i1}	β_{i2}	β_{i1}	β_{i2}
Australia	0.47 (0.05)	−0.29 (0.36)	0.53 (0.23)	2.13 (1.65)	0.41 (0.22)	0.03 (1.91)
Austria	−1.39 (0.23)	−0.80 (0.36)	−2.37 (0.74)	0.01 (1.16)	−1.57 (0.75)	−0.10 (1.35)
Belgium	−0.85 (0.16)	0.29 (0.21)	−0.71 (0.85)	0.06 (1.18)	−0.95 (0.53)	−0.06 (0.87)
Canada	0.47 (0.08)	0.26 (0.31)	0.84 (0.21)	1.36 (0.78)	0.60 (0.20)	0.97 (0.82)
Denmark	−0.30 (0.14)	2.06 (0.43)	−0.42 (0.47)	4.54 (1.41)	−0.37 (0.39)	2.45 (1.36)
Finland	0.09 (0.07)	0.67 (0.25)	0.03 (0.26)	1.41 (0.85)	−0.12 (0.26)	1.28 (0.83)
France	1.64 (0.42)	−0.47 (0.20)	4.59 (1.10)	−0.72 (0.52)	2.52 (0.98)	−0.69 (0.47)
Germany	−0.30 (0.15)	1.27 (0.25)	−0.29 (0.34)	1.27 (0.56)	−0.39 (0.34)	1.25 (0.60)
Great Britain	−0.56 (0.11)	−0.58 (0.21)	−0.56 (0.40)	−0.47 (0.76)	−0.50 (0.38)	−0.43 (0.79)
Greece	1.04 (0.04)	0.44 (0.26)	1.35 (0.22)	2.41 (1.28)	1.11 (0.27)	0.77 (1.61)
Italy	0.92 (0.08)	−0.14 (0.19)	1.41 (0.33)	1.15 (0.76)	0.61 (0.43)	−0.49 (1.70)
Japan	−3.16 (0.51)	−1.28 (0.34)	−3.38 (1.66)	−1.75 (1.11)	−3.15 (3.27)	−1.34 (1.94)
Korea	0.23 (0.10)	0.01 (0.14)	0.10 (0.18)	0.27 (0.26)	0.16 (0.15)	0.07 (0.21)
Netherlands	−0.06 (0.25)	1.28 (0.29)	0.07 (0.63)	2.21 (0.75)	−0.23 (0.41)	2.11 (0.53)
Norway	−0.68 (0.20)	1.24 (0.22)	−1.81 (0.67)	2.50 (0.73)	−1.48 (0.69)	2.07 (0.74)
Spain	0.23 (0.14)	−1.83 (0.63)	0.30 (0.40)	−1.06 (1.75)	0.25 (0.68)	−1.25 (3.19)
Sweden	−0.03 (0.36)	−1.26 (0.24)	0.11 (0.89)	−2.01 (0.60)	−0.38 (0.73)	−1.70 (0.64)
Switzerland	1.02 (0.25)	1.59 (0.32)	1.46 (1.01)	1.28 (1.20)	0.85 (1.21)	1.33 (1.94)

Note: Standard errors are given in parentheses.

Table 2: Additional country-by-country estimates of the cointegrating relationship, $e_{it} = \beta_{i0} + \beta_{i1}(m_{it}^{*} - m_{it}) + \beta_{i2}(y_{it}^{*} - y_{it})$

(1)	(2)	(3)	(4)	(5)	(6)	(7)
	JOH-ML estimates		Unrestricted Dynamic SUR estimates		ARDL estimates, common time effect	
Country	β_{i1}	β_{i2}	β_{i1}	β_{i2}	β_{i1}	β_{i2}
Australia	0.51 (0.11)	1.32 (1.09)	0.23 (0.06)	0.18 (0.18)	0.77 (0.26)	−1.75 (1.46)
Austria	−0.61 (0.90)	−6.24 (2.77)	−0.86 (0.12)	−0.24 (0.12)	1.47 (0.26)	−0.44 (0.85)
Belgium	−1.53 (0.30)	−1.93 (0.57)	−0.55 (0.09)	0.12 (0.16)	0.57 (1.94)	−1.06 (5.84)
Canada	0.82 (0.16)	1.52 (0.74)	0.04 (0.05)	0.20 (0.13)	−1.44 (1.29)	1.85 (1.68)
Denmark	−1.96 (0.98)	11.46 (4.71)	0.68 (0.13)	0.06 (0.18)	−0.04 (0.60)	−3.45 (2.06)
Finland	−0.34 (0.16)	2.07 (0.51)	0.05 (0.06)	−0.08 (0.18)	−0.08 (0.16)	−0.61 (0.37)
France	3.62 (0.72)	−0.50 (0.33)	−1.91 (0.25)	−0.30 (0.17)	0.52 (0.50)	−0.28 (0.92)
Germany	−0.66 (0.16)	1.12 (0.27)	−0.00 (0.07)	0.03 (0.14)	1.28 (1.05)	−2.02 (3.78)
Great Britain	−0.36 (0.20)	−1.10 (0.46)	−0.80 (0.11)	−0.20 (0.20)	−0.37 (0.14)	1.44 (0.53)
Greece	1.33 (0.16)	2.88 (1.10)	0.94 (0.04)	−0.37 (0.17)	1.21 (0.16)	0.54 (0.81)
Italy	0.94 (0.53)	6.99 (4.44)	0.82 (0.09)	−0.29 (0.13)	0.55 (0.39)	−1.16 (0.62)
Japan	14.93 (9.79)	−3.41 (1.70)	−1.43 (0.32)	−0.76 (0.21)	2.19 (0.41)	−1.64 (0.78)
Korea	−0.49 (0.37)	1.02 (0.56)	0.08 (0.07)	0.10 (0.09)	0.64 (0.37)	−0.44 (0.42)
Netherlands	−0.33 (0.22)	2.34 (0.27)	−0.05 (0.11)	0.55 (0.16)	1.04 (0.27)	−0.83 (0.83)
Norway	−2.54 (0.45)	3.18 (0.47)	−0.27 (0.13)	0.69 (0.12)	0.74 (0.21)	0.05 (0.05)
Spain	2.62 (1.54)	10.67 (7.63)	0.15 (0.10)	−0.72 (0.31)	0.41 (0.65)	−0.68 (1.44)
Sweden	−0.92 (0.38)	−3.70 (0.38)	−0.46 (0.25)	−0.43 (0.17)	−0.27 (0.16)	−0.60 (0.28)
Switzerland	−1.74 (1.55)	−4.07 (2.82)	0.84 (0.18)	0.24 (0.21)	1.32 (0.43)	−2.07 (2.50)

Note: Standard errors are given in parentheses.

Table 3: Country-by-country cointegration test results

(1) Country	(2) Augmented Engle-Granger (1987) test[a]	(3) Johansen (1991) trace trace[b]	(4) Degrees-of-freedom corrected trace test
Australia	−2.16 (3)	35.30* (7)	25.45
Austria	−2.69 (4)	30.19* (8)	20.60
Belgium	−2.29 (3)	33.34* (8)	22.75
Canada	−2.99 (6)	45.30** (5)	36.03**
Denmark	−2.48 (0)	36.73** (3)	31.83*
Finland	−3.14 (6)	26.43 (5)	21.02
France	−2.90 (4)	27.29† (4)	22.69
Germany	−2.67 (4)	38.39** (7)	27.68†
Great Britain	−3.20 (7)	25.91 (7)	18.68
Greece	−2.49 (4)	22.56 (6)	17.11
Italy	−2.63 (4)	30.40* (4)	25.28
Japan	−1.68 (5)	41.42** (8)	28.26†
Korea	−1.19 (3)	21.78 (6)	16.52
Netherlands	−2.50 (4)	60.91** (6)	46.21**
Norway	−2.96 (7)	30.83* (7)	22.23
Spain	−2.09 (8)	27.09† (1)	25.32
Sweden	−2.65 (5)	51.76** (8)	35.32*
Switzerland	−2.37 (4)	27.01 (4)	22.46

Notes: †,*,** indicate significance at the 10, 5, and 1 percent levels, respectively; endogenously selected lag orders for the different tests are given in parentheses.
[a]One-sided (lower-tail) test of the null hypothesis that e_t, $m_t^* - m_t$, and $y_t^* - y_t$ are not cointegrated; Phillips and Ouliaris (1990) 10, 5, and 1 percent critical values equal −3.45, −3.77, and −4.31, respectively.
[b]One-sided (upper-tail) test of the null hypothesis that e_t, $m_t^* - m_t$, and $y_t^* - y_t$ are not cointegrated; Osterwald-Lenum (1992) 10, 5, and 1 percent critical values equal 26.79, 29.68, and 35.65, respectively.

Table 4: Panel estimates of the cointegrating relationship, $e_{it} = \beta_{i0} + \beta_1(m^*_{it} - m_{it}) + \beta_2(y^*_{it} - y_{it})$

(1)	(2)	(3)	(4)	(5)	(6)	(7)
	LSDV estimates		Bias-adjusted LSDV estimates		Restricted dynamic SUR estimates	
Panel	β_1	β_2	β_1	β_2	β_1	β_2
EC[a] (9)	0.78 (0.03)	−0.20 (0.10)	0.84 (0.14)	−0.18 (0.19)	0.64 (0.04)	−0.36 (0.15)
EMS[b] (6)	0.44 (0.06)	0.16 (0.11)	0.51 (0.21)	0.21 (0.10)	0.13 (0.05)	−0.19 (0.12)
Group of 6[c]	0.33 (0.07)	−0.61 (0.12)	0.40 (0.22)	−0.69 (0.22)	0.46 (0.10)	−0.73 (0.20)
Group of 10[d]	0.30 (0.06)	−0.28 (0.09)	0.38 (0.18)	−0.30 (0.17)	0.16 (0.04)	−0.22 (0.08)
All 18	0.63 (0.02)	−0.51 (0.04)	0.68 (0.09)	−0.56 (0.13)	0.60 (0.01)	−0.47 (0.02)
	LSDV estimates, common time effect		Bias-adjusted LSDV estimates, common time effect		Pooled mean group estimates, common time effect	
Panel	β_1	β_2	β_1	β_2	β_1	β_2
EC[a] (9)	0.86 (0.02)	−0.62 (0.08)	0.90 (0.08)	−0.73 (0.14)	0.39 (0.12)	−1.43 (0.32)
EMS[b] (6)	0.80 (0.03)	−0.45 (0.08)	0.86 (0.11)	−0.54 (0.12)	0.11 (0.17)	−0.32 (0.33)
Group of 6[c]	0.70 (0.05)	−1.43 (0.11)	0.75 (0.21)	−1.43 (0.11)	0.48 (0.30)	−1.29 (0.69)
Group of 10[d]	0.73 (0.05)	−1.15 (0.09)	0.80 (0.16)	−1.35 (0.16)	0.68 (0.20)	−2.74 (0.53)
All 18	0.86 (0.02)	−0.76 (0.03)	0.92 (0.07)	−0.85 (0.09)	0.90 (0.06)	−0.73 (0.09)

Notes: Standard errors are given in parentheses.
[a]Belgium, Denmark, France, Germany, Great Britain, Greece, Italy, Netherlands, Spain.
[b]Belgium, Denmark, France, Germany, Italy, Netherlands.
[c]Canada, France, Germany, Great Britain, Italy, Japan.
[d]Belgium, Canada, France, Germany, Great Britain, Italy, Japan, Netherlands, Sweden, Switzerland.

Table 5: Sample standard deviations for nominal exchange rates, relative money supplies, and relative real output levels

(1)	(2)	(3)	(4)	(5)	(6)	(7)
	Nominal exchange rate		Relative money supply		Relative real output level	
Panel	Country-by-country average stan. deviation	Panel stan. deviation	Country-by-country average stan. deviation	Panel stan. deviation	Country-by-country average stan. deviation	Panel stan. deviation
EC[a] (9)	0.282	2.328	0.221	2.742	0.085	0.091
EMS[b] (6)	0.221	2.237	0.133	2.482	0.083	0.089
Group of 6[c]	0.225	2.804	0.130	3.725	0.081	0.102
Group of 10[d]	0.225	2.332	0.113	3.069	0.081	0.097
All 18	0.249	2.233	0.222	2.680	0.108	0.182

Notes: Standard errors are given in parentheses.
[a]Belgium, Denmark, France, Germany, Great Britain, Greece, Italy, Netherlands, Spain.
[b]Belgium, Denmark, France, Germany, Italy, Netherlands.
[c]Canada, France, Germany, Great Britain, Italy, Japan.
[d]Belgium, Canada, France, Germany, Great Britain, Italy, Japan, Netherlands, Sweden, Switzerland.

Table 6: Panel cointegration test results

(1)	(2)	(3)	(4)	(5)	(6)
Panel	PC [e]	DF_{ρ}^{*} [e]	DF_{t}^{*} [e]	SURPADF [e]	MADF [f]
EC[a] (9)	-12.47^{*}	-9.28^{*}	-2.14^{*}	$-4.97^{\dagger}$ (4)	47.63^{*} (4)
EMS[b] (6)	-9.11	-7.72	-1.75	-4.02 (5)	22.20 (5)
Group of 6[c]	-19.66^{**}	-16.84^{*}	-2.81^{*}	-4.84^{*} (5)	$28.06^{\dagger}$ (5)
Group of 10[d]	-14.71^{*}	-12.77^{**}	-2.62^{**}	-5.87^{**} (5)	39.75 (5)
All 18	-17.23^{*}	-11.51^{**}	-2.73^{*}	-8.05^{**} (4)	93.42^{*} (5)

Notes: †,*,** indicate significance at the 10, 5, and 1 percent levels, respectively.
[a]Belgium, Denmark, France, Germany, Great Britain, Greece, Italy, Netherlands, Spain.
[b]Belgium, Denmark, France, Germany, Italy, Netherlands.
[c]Canada, France, Germany, Great Britain, Italy, Japan.
[d]Belgium, Canada, France, Germany, Great Britain, Italy, Japan, Netherlands, Sweden, Switzerland.
[e]One-sided (lower-tail) test of the null hypothesis that e_t, $m_t^* - m_t$, and $y_t^* - y_t$ are not cointegrated; significance is based on critical values generated using a nonparametric version of Groen (2000) bootstrap procedure described in the text; endogenously selected lag orders are given in parentheses.
[f]One-sided (upper-tail) test of the null hypothesis that e_t, $m_t^* - m_t$, and $y_t^* - y_t$ are not cointegrated; significance is based on critical values generated using a nonparametric version of Groen (2000) bootstrap procedure described in the text; endogenously selected lag orders are given in parentheses.

Table 7: Test results for the null hypothesis that $\beta_{i1} = \beta_1$ and $\beta_{i2} = \beta_2$ for $i = 1, \dots, N$

(1)	(2)	(3)	(4)	(5)
	Mark, Ogaki, and Sul (2000) homogeneity test		Pesaran, Shin, and Smith (1999) homogeneity test	
Countries	Wald statistic	p-value	Likelihood-ratio statistic	p-value
EC[a] (9)	109.44	0.00	33.96	0.01
EMS[b] (6)	22.01	0.02	14.52	0.15
Group of 6[c]	108.39	0.00	18.35	0.05
Group of 10[d]	248.71	0.00	25.67	0.11
All 18	1315.07	0.00	88.50	0.00

Notes: [a]Belgium, Denmark, Germany, Great Britain, Greece, Italy, Netherlands, Spain.
[b]Belgium, Denamrk, France,Germany, Italy, Netherlands.
[c]Canada, France, Germany, Great Britain, Italy, Japan.
[d]Belgium, Canada, France, Germany, Great Britain, Italy, Japan, Netherlands, Sweden, Switzerland.

Table 8: Comparison of out-of-sample forecast performance

(1)	(2)	(3)	(4)	(5)	(6)
Heterogeneous Estimator	1-step-ahead	4-step-ahead	8-step-ahead	12-step-ahead	16-step-ahead
OLS	1.008	1.005	0.936	0.881	0.852
FM-OLS	1.016	1.035	0.975	0.909	0.907
DOLS	1.010	1.011	0.936	0.873	0.839
JOH-ML	1.018	1.042	0.950	0.855	0.830
UDSUR[a]	1.007	0.986	0.907	0.865	0.846
ARDL	1.010	1.013	1.017	1.001	0.966

Notes: Table entries report the ratio $RMSE_2/RMSE_1$, where $RMSE_1$ is the root-mean-squared prediction error for the predictive regression based on the heterogeneous estimates and $RMSE_2$ is the root-mean-squared prediction error for the predictive regression based on the homogeneous estimates.
[a]Unrestricted dynamic SUR estimator.

Table 9: Summary statistics for the LSDV pooled estimates that allow for a common time effect when the true data generating process is characterized by heterogeneity

(1) Panel	(2) Mean	(3) Sample stan. dev.	(4) Maximum	(5) Minimum	(6) Median
A. LSDV β_1 estimate					
EC[a] (9)	0.782	0.190	1.674	0.006	0.784
EMS[b] (6)	0.494	0.378	2.024	−0.911	0.523
Group of 6[c]	0.572	0.476	2.579	−1.824	0.609
Group of 10[d]	0.477	0.387	1.567	−1.432	0.541
All 18	0.824	0.297	2.576	−0.115	0.802
B. LSDV β_2 estimate					
EC[a] (9)	−0.033	0.416	1.155	−1.417	−0.028
EMS[b] (6)	0.006	0.375	1.38	−1.393	0.021
Group of 6[c]	−0.158	0.576	2.006	−2.721	−0.126
Group of 10[d]	−0.152	0.492	1.400	−2.270	−0.161
All 18	0.282	0.867	4.287	−1.633	0.155

Notes: [a]Belgium, Denmark, Germany, Great Britain, Greece, Italy, Netherlands, Spain.
[b]Belgium, Denamrk, France,Germany, Italy, Netherlands.
[c]Canada, France, Germany, Great Britain, Italy, Japan.
[d]Belgium, Canada, France, Germany, Great Britain, Italy, Japan, Netherlands, Sweden, Switzerland.