# PANEL DATA EVIDENCE ON THE DEMAND FOR MONEY*

## 1. Introduction

Nearly all modern macroeconomic theories are built upon the behavior of the money market and the demand for money. Economists and policy makers alike have delved into the subject matter in order to understand the surrounding implications for monetary policy. In particular, researchers have tried to address and empirically validate two specific theoretical models. The first, the quantity theory demand for money, asserts that the income elasticity of real money balances is equal to one and that interest rates have no effect on the demand for real money balances (i.e. $\eta_Y = 1$ and $\eta_R = 0$, where $\eta_z$ denotes the elasticity of the demand for real balances with respect to $z$). The second, the Baumol (1952) and Tobin (1956) transactions theory of the demand for money, asserts that agents are preoccupied by the decision of when and how often to exchange bonds for money. Ultimately, the testable hypotheses are that a rise in real income leads to a less-than-proportionate increase in the average holding of real money and that interest rates do have a negative role in the demand for money (in particular, $\eta_Y = 1/2$ and $\eta_R = -1/2$).

Within this vast literature, contemporary empirical studies have focused on estimating the conventional money demand function by means of error-correction models – see, for example, Sriram (1999) and Serletis (2007) for a brief review and references. However, these models only emphasize the time series characteristics of the data. As a result, the elasticity estimates derived from the time-series approach seem to be sensitive to the univariate and multivariate time series properties of the money demand variables, the sample period, the functional form, and the definition of the underlying variables. Others, such as Fisher and Seater (1993) and King and Watson (1997) have

* Serletis gratefully acknowledges financial support from the Social Sciences and Humanities Research Council of Canada (SSHRCC).

tried to develop more solid procedures in order to test the long-run neutrality and superneutrality hypotheses. Serletis and Koustas (1998, 2001), for example, have applied both methodologies and have found evidence that money is neutral in the long-run for most countries. Yet, when international comparisons are made applied researchers have typically neglected the cross-section component of the data.

More recently, Fujiki, Hsiao, and Shen (2002) and Fischer (2005) have estimated cross-regional money demand in both Japan and Switzerland using traditional panel data methodology. Moreover, Serletis and Vaccaro (2006) have used cross-country data (for 48 countries over the 1980-1995 period) to investigate the long-run relationship between (both narrow and broad) money and interest rates, real GDP, institutions, financial structure, and financial development. They have shown that the interest and income elasticities of real money balances are fairly stable and conform to the theoretical predictions of the quantity theory demand for money. As well, they have found that institutions, financial structure, and financial development do play a role in the demand for money in an aggregate setting, albeit a limited role. However, Serletis and Vaccaro (2006) have also shown that the assumption that all of the countries can be treated as a homogeneous unit can cause systematic distortions. Specifically, they utilized unsupervised Bayesian methods, based on finite mixture models and mathematical properties, to cluster the data set into two distinct groups. Regressions based on each of the partitioned data sets displayed heterogeneity with respect to the influence institutions, financial structure, and financial development have on money demand, for each of the two groups.

It seems that a panel data approach to investigating money demand issues is advantageous, because it allows the researcher to sort out economic effects that may not be distinguishable with the use of either cross-section or time series data alone. In the case of international comparisons, the panel consists of countries overtime. As such, through this methodology we can enhance our econometric modeling and hypotheses testing by investigating possible heterogeneity across these units. In particular, the techniques of panel data estimation can take such heterogeneity into account by allowing for country-specific variables. Secondly, panel data analysis minimizes the bias that might result if we aggregate countries into a broad homogeneous unit and simply use the cross-section. Lastly, by combining the time series of cross-sectional observations, panel data gives us more informative data, variability, degrees of freedom, less collinearity among variables, and added efficiency – see, for example, Gujarati (2003, Chapter 16) for an extensive discussion.

In addition, within the contemporary econometrics of panel data, there have been extensive interest and contributions made to nonstationary panel time-series models and dynamic panel data – see Banerjee (1999), Phillips and Moon (2000), and Baltagi and Kao (2000) for concise surveys and references. As Baltagi and Kao (2000, p. 8) point out,

> "the hope of the econometrics of nonstationary panel data is to combine the best of both worlds: the method of dealing with nonstationary data from the time series and the increased data and power from the cross-section. The addition of the cross-section dimension, under certain assumptions, can act as repeated draws from the same distribution. Thus as the time and cross-section dimension increase, panel statistics can be derived which converge in distribution to normally distributed random variables".

Two strands of the literature mentioned above, which have not been applied to money demand research, are testing for panel unit roots and panel cointegration. Regarding panel unit roots, there have been many recent contributions. Notable advances have been made by Levin, Lin, and Chu (2002), Breitung (2000), Im, Pesaran, and Shin (2003), Maddala and Wu (1999), Choi (2001), Sarno and Taylor (1998), and Hardi (2000), among others. The tests derived by these authors are quite diverse with respect to their construction and interpretation, and have been widely used in the modern purchasing power parity, growth and convergence, real GDP, and exchange rate literatures. Regarding panel cointegration, notable advances have been made by McCoskey and Kao (1998), Kao (1999), and Pedroni (1999, 2000, 2001, 2004), among others. These panel cointegration and panel cointegrating vector tests are able to extensively accommodate heterogeneous dynamics across individual members of the panel and have allowed researchers to directly test the long-run equilibrium relationships highlighted above. Banerjee (1999) summarizes these recent developments in both subfields quite nicely as

> "in other instances where a new literature comes to be seen to be significant, the aggregate has turned out to be greater than the sum of its parts and the theory and practice of integrated series in panel data have given rise to a set of interesting and surprising results which are uniquely its own".

In this paper, our main objective is to embark on the first preliminary investigation of cross-country money demand which exploits panel routines to investigate the long-run relationship between real money balances, nominal interest rates, and real income. Our research utilizes traditional panel methodology as well as recently developed state-of-the art panel unit root and panel cointegration techniques, in order to test diverse aggregate long-run theories of money demand and examine possible heterogeneity within the dataset. In our opinion, a 48 multi-country setting over the 1980-95 time period provides itself particularly well to both cross-country and aggregate comparisons of the issue under discussion.

The organization of this paper is as follows. In Section 2 we describe the data and the underlying sources of collection and origin. Section 3 outlines the conventional panel data methodology utilized to estimate both narrow and broad money demand functions and presents estimates using traditional panel data models. Section 4 summarizes the state-of-the-art contemporary panel routines used to investigate the aggregate group relationships between real monetary aggregates, interest rates and real gross domestic product. In the same section we display the results of these new innovative panel time-series procedures applied to the demand for money. The final section concludes the paper and outlines the implications of our findings.

## 2. The data

The narrow definition of money chosen is what we shall refer to as M1. The International Monetary Fund (IMF) and standard monetary textbooks define such a narrow measure as, transferable deposits (demand deposits) and currency outside of banks. The broad definition of money chosen is what we shall refer to as M2. This broad measure is identified as M1 plus quasi money (time, savings, and foreign currency deposits). For the 48 countries included in the study – see Table 1 – annual data pertaining to both measures were collected over the 1980-1995 period from the IMF International Financial Statistics (IFS), the World Development Indicators (WDI), and various central banks in local currency units – this data set is also part of the Serletis and Vaccaro (2006) study. The data were then converted to United States dollars by using the U.S. dollar per local currency unit 1995 average exchange rate for each country.

In order to analyze the monetary aggregates described above in real terms, we then collected data from the WDI on the consumer price index (CPI) for each country with a base year of 1995. The average was then taken to obtain a single observation for each country.

Each of the monetary aggregates was then deflated by the consumer price index for each of the 48 countries to obtain a real measure. Although the GDP deflator would have been the ideal price index to use, it was not exploited due to data availability and base year issues. However, for those countries for which we found both, a comparison was made and differences between the two indices were minor if not nil. At any rate, the CPI is the most publicly reported price index. Constant 1995 U.S. dollar GDP data were also collected from the WDI for each country.

TABLE 1

*Countries*

| | |
|---|---|
| Argentina | Kenya |
| Australia | Malaysia |
| Austria | Mexico |
| Belgium | Netherlands |
| Brazil | New Zealand |
| Canada | Norway |
| Chile | Pakistan |
| Colombia | Panama |
| Cyprus | Peru |
| Denmark | Philippines |
| Ecuador | Portugal |
| Egypt | South Africa |
| Finland | Spain |
| France | Sri Lanka |
| Germany | Sweden |
| Ghana | Switzerland |
| Greece | Taiwan, China |
| Honduras | Thailand |
| India | Trinidad and Tobago |
| Ireland | Turkey |
| Israel | Tunisia |
| Italy | United Kingdom |
| Jamaica | United States |
| Japan | Zimbabwe |

With regards to short term interest rate data, there were some data availability issues. We could not find a uniformly defined interest rate series for all 48 countries. As a result, data were first collected for countries for which there existed a 90-day treasury bill rate or the local equivalent. Subsequently, data were collected for those countries for which there existed a money market rate. For those countries which neither existed, a deposit rate was collected. Collecting interest rate data from Latin and South American countries in some cases was quite tedious.

## 3. Traditional panel specifications and results

Panel data refers to data for $N$ different entities observed over $T$ different time periods. A panel data set is advantageous because it allows us to sort out economic effects that may not be distinguishable with the use of either cross-section or time series data alone. In this section, we outline the traditional panel data methodology and use it to estimate conventional money demand functions for both narrow and broad specifications. It is to be noted that we are dealing with an unbalanced panel, as not all of the countries in our study have the same number of time series observations.

We begin our study by using the three most common estimators: pooled ordinary least squares (OLS), the fixed effects model (FEM), and the random effects model (REM). The assumptions of these estimators do differ and each has its own drawbacks, as will be discussed shortly. The pooled model can be expressed as

$$m_{it} = b_1 + b_2 R_{it} + b_3 Y_{it} + \varepsilon_{it} \tag{1}$$

where $i$ represents country, with $i=1,....48$, and $t$ denotes time, with $t=1,...,15$. $R$ and $Y$ denote the natural logarithm of the opportunity cost and transactions variables, respectively. For the pooled model it is assumed that $E(\varepsilon_{it})=0$ for all $i$ and $t$, $E(\varepsilon_{it}^2)=\sigma^2_\varepsilon$ and $E(\varepsilon_{it}, \varepsilon_{js})=0$ for all $s\neq t$ or all $i\neq j$. The major pitfall of the pooled model is that it ignores heterogeneity across countries with respect to unobservable characteristics, either for lack of variation or as a deliberate modeling choice. Hsiao (2003) points out that either reason may cause the pooled estimator to be biased. Hence, we use it as a base specification to make comparisons against. A total of $NT-3$ degrees of freedom would be involved for this estimator.

In general, the most common procedures to account for heterogeneity in panel data are the FEM and REM estimators. Between the two estimators, they can account for heterogeneity across units, by means of decomposing the effects of unobservable factors into effects specific to cross-sectional units, to time-periods, and to both cross-sectional units and time-periods – see Hsiao (2003, p. 97) for an introduction and comprehensive discussion of the two estimators. The fixed effects model which we are interested in can be expressed as,

$$m_{it} = b_1 + \alpha_1 D_{2i} + \alpha_2 D_{3i} + \cdots + \alpha_{47} D_{48i}\ b_2 R_{it} + b_3 Y_{it} + \varepsilon_{it} \tag{2}$$

where, the $\alpha$'s are the differential intercept coefficients representing a time-invariant group specific attribute. The $D$'s represent country specific dummy variables. By allowing the intercept to vary, we can investigate whether or not country specific attributes shift the money demand function. However, as a group we should observe a long-run money demand theory, such as the quantity theory demand for money or the Baumol-Tobin transactions theory, to hold.

It must be noted that the inclusion of dummies does not directly identify the sources which may cause the intercept to shift over countries. However, the cross-country estimates obtained by Serletis and Vaccaro (2006) can give some kind of idea and intuition to which institution, financial development, and financial service variables may be useful for future panel modeling, once data availability and collection issues are overcome. In addition, another pitfall of the FEM is that a substantial number of degrees of freedom are lost with the addition of so many coefficients. For example, in our case where we allow the constant to vary, there would be 47 dummy differential intercept coefficients, an intercept coefficient for the base country, and 2 coefficients for the interest rate and real income elasticities. Clearly, specification and diagnostic tests would have to be conducted to determine which of the three models is preferred.

Alternatively, the REM model reflects the lack of knowledge about the model through the disturbance term. The REM does so by using a pooled cross-section and time series model in which error terms may be correlated across time and individual units. The REM model can be expressed as,

$$m_{it} = b_1 + b_2 R_{it} + b_3 Y_{it} + \varepsilon_{it} \tag{3}$$

$$\varepsilon_{it} = u_i + v_t + w_{it} \tag{4}$$

where $u_i \mathrm{b} N(0, \sigma_u^2)$ represents the cross-section error component, $v_t \mathrm{b} N(0, \sigma_v^2)$ signifies the time series component and $w_{it} \mathrm{b} N(0, \sigma_w^2)$ denotes the combined error component. It is assumed that individual error components are uncorrelated with each other and are not autocorrelated across both cross-section and time series. At the same time the error term would consist of three components and would have variance

$$\text{Var } (\varepsilon_{it}) = \sigma_u^2 + \sigma_v^2 + \sigma_w^2 \tag{5}$$

If both $\sigma_u^2$ and $\sigma_v^2$ are equal to 0, the error term consists of a single combined white noise disturbance and the pooled model is preferred. When the combined error component $\sigma_w^2$ equals zero, then the fixed effects model is preferred. The REM is estimated as a two-stage generalized least-squares regression. Typically the REM is considered an intermediate model which lies between the extreme of a zero combined error component (FEM) and an infinitely large combined error component (pooled model).

In order to obtain estimates of the traditional panel models that we just described, we utilize the EViews 5 quantitative micro software. To begin, we present the panel descriptive statistics for the conventional money demand variables in Table 2. There is a wide variation of logged real money supplies, logged nominal interest rates, and logged real outputs across the panel. Table 3 displays the

TABLE 2 - *Panel Descriptive Statistics: 1980-1995*

| Series | Mean | Minimum | Maximum | Standard deviation | Cross sections, $N$ | N×T |
|---|---|---|---|---|---|---|
| log M1 | 18.620 | 14.828 | 23.632 | 1.991 | 48 | 764 |
| log M2 | 19.787 | 15.046 | 24.795 | 2.046 | 48 | 762 |
| log $R$ | 2.617 | 0.190 | 16.087 | 1.184 | 48 | 752 |
| log $Y$ | 25.229 | 21.685 | 29.624 | 1.849 | 48 | 768 |

TABLE 3 - *Money Demand: Conventional Panel Data Estimators*

| | Model | | | | | |
|---|---|---|---|---|---|---|
| | Pooled OLS | | Fixed Effects[b] | | Random Effects | |
| | Dependent Variable[a] | | | | | |
| Regressors | M1 | M2 | M1 | M2 | M1 | M2 |
| Constant | -6.559** | -6.617** | -3.506** | -12.655** | -5.096** | -9.811** |
| | (0.332) | (0.272) | (1.151) | (0.966) | (0.846) | (0.685) |
| log $R$ | -0.114** | -0.145** | -0.015 | 0.031** | -0.016 | -0.024** |
| | (0.020) | (0.016) | (0.011) | (0.009) | (0.011) | (0.009) |
| log $Y$ | 1.009** | 1.061** | 0.878** | 1.282** | 0.941** | 1.170** |
| | (0.012) | (0.010) | (0.045) | (0.038) | (0.033) | (0.026) |
| $\bar{R}^2$ | 0.891 | 0.931 | 0.986 | 0.991 | 0.499 | 0.705 |
| $F$ | 3081.5** | 5087.6** | 1118.9** | 1687.2** | 374.2** | 893.5** |
| DW | 0.051 | 0.084 | 0.322 | 0.417 | 0.300 | 0.373 |
| LR[c] | | | 1609.6** | 1567.2** | | |
| Hausman[d] | | | | | 6.978* | 30.197** |

*Notes*:

[a] Both aggregates are in real terms and logged. Standard errors are given in parentheses, and * and ** indicate significance at the 5 and 1% levels, respectively. The Swamy and Arora algorithm is used to estimate the component variances for the REM.

[b] The FEM only allows the constant to vary. The estimated coefficients of the group-specific effects are omitted.

[c] The LR statistic refers to a test of the null hypothesis of the pooled cross-section model against the fixed effects model. The statistic has a $\chi^2$ distribution with ($N$–1) degrees of freedom, where $N$ is the number of cross-section units. Note that the estimates of the FEM include coefficients for group-specific effects.

[d] The Hausman test is a test with the null hypothesis of the random effects model against the fixed effects model. The statistic has a $\chi^2$ distribution with 2 degrees of freedom.

results for the estimates of (1)-(3) for each monetary aggregate as the dependent variable. In general, the three models are either in accordance with the quantity theory demand for money, or come close for the group of 48 countries. In the pooled model, we test the restriction that the income coefficient is equal to one and we cannot reject the restriction placed on the real income elasticity, for both aggregates. The interest elasticity is negative and significant, but closer to the quantity theory than the Baumol-Tobin prediction.

The next two columns show the fixed effects estimates. This specification allows the constant to vary according to country specific attributes. However, although we do not report the country specific results, the FEM displayed heterogeneity in the constant with the majority of the country specific coefficients being significant. For example, the country specific constants varied from -2.129 for Belgium to 1.232 for Japan, when we treated M1 as the dependent variable. When we treated M2 as the dependent variable, the country specific constants varied from -2.291 for Belgium to 1.326 for Cyprus. The estimated interest rate elasticity is not significant with M1 as the dependent variable. However, with M2 as the dependent variable it becomes marginally positive and significant. Regarding the real income elasticity, the estimates varied from 0.878 to 1.282 for each of the real monetary aggregates.

The last two columns report the results for the REM. In contrast to the fixed effects estimates, the random effects estimates are more supportive of the quantity theory demand for money. The interest elasticity is not significant at conventional levels for either aggregate and the income elasticity varies from 0.94 for M1 to 1.17 for M2.

In order to make selection between the models we experimented with, we conducted a variety of specification tests. The likelihood ratio test statistic (LR) is highly significant in both cases. This allows us to reject the null hypothesis of the pooled model over the fixed effects model, for each monetary aggregate. The Hausman test statistic is also highly significant, indicating that the fixed effects model is preferred over the random effects model. In summary, these standard panel specification tests show that the fixed effects model is to be preferred over the pooled cross-section and the random effects models. This conclusion indicates that there is a great deal of heterogeneity among the countries.

However, inspection of the Durbin-Watson (DW) statistics indicates significant residual serial correlation and specification errors. In particular, the DW statistics are very low for all of the models we experimented with. Therefore, we conclude that these estimators are biased and may not be consistent and that further hypothesis testing within these models would be spurious. As such, we next turn to recent, state-of-the-art developments in panel data estimation in order to further our investigation of money demand issues.

## 4. Contemporary panel methodology

The panel unit root tests that we consider are from Maddala and Wu (1999), Breitung (2000), Hardi (2000), Levin, Lin, and Chu (2002), and Im, Pesaran, and Shin (2003). Regarding panel cointegration, we will exploit Pedroni's (1999, 2000, 2001, 2004) recent contributions – although we would have also liked to include the Taylor and Sarno (1998) and Sarno and Taylor (1998) MADF test based on SUR, we did not, because we are dealing with an unbalanced panel; in order to apply their methodology a balanced panel is required.

### *4.1 Panel Unit Root Tests*

Although the panel unit root tests are similar, they are not identical, and as such we begin by briefly outlining the various tests through the following AR(1) process for panel data

$$y_{it} = \rho_i y_{it-1} + \boldsymbol{X'}_{it}\delta + \varepsilon_{it}$$

where, as before, $i = 1,2,\ldots,N$ cross-section units that are observed over periods $t = 1,2,\ldots,T$. The exogenous variables in the model, which may include any fixed effects or individual trends, are represented by $\boldsymbol{X}$, the autoregressive coefficients are denoted by $\rho_i$, and the errors, $\varepsilon_{it}$, are assumed to be mutually i.i.d. If $|\rho_i|<1$, then $y_i$ is said to be weakly (trend-) stationary. However, if $|\rho_i|=1$ then $y_i$ has a unit root. With regards to testing, two assumptions can be made about $\rho_i$. Tests which are considered first generation tests, assume that the persistence parameters are common across cross-sections, so that $\rho_i = \rho$, for all $i$. The Levin, Lin, and Chu (2002), Breitung (2000), and Hardi (2000) tests make this assumption. Alternatively, second generation tests allow $\rho_i$ to vary freely across the cross-section units. The Im, Pesaran, and Shin (2003) and Maddala and Wu (1999) tests are of this form.

Levin, Lin, and Chu (2002) and Breitung (2000) both consider the following basic ADF specification

$$\Delta y_{it} = \alpha y_{it-1} + \sum_{j=1}^{p_i} \beta_{ij} \Delta y_{it-j} + X'_{it} \delta + \varepsilon_{it}$$

where a common $\alpha = \rho - 1$ is assumed, but the lag order for the difference terms, $p_i$, is allowed to vary across the units. Under the null hypothesis, $H_0 : \alpha = 0$, there is a unit root for all series, whereas under the alternative, $H_A : \alpha < 0$, none of the series contain a unit root.

The methodology described in Levin, Lin, and Chu (2002) develops estimates of $\alpha$ from proxies for $\Delta y_{it}$ and $y_{it}$, which are standardized and free of autocorrelations and deterministic components. For a given set of lag orders supplied by the researcher, the Levin, Lin, and Chu (2002) algorithm begins by estimating two additional sets of equations. First, both $\Delta y_{it}$ and $y_{it-1}$ are regressed on the lag terms $\Delta y_{it-j}$ (for $j = 1,2,..., p_i$) and then on the exogenous variables $\boldsymbol{X_{it}}$. The estimated coefficients from these two regressions can be denoted as $(\hat{\boldsymbol{\beta}}, \hat{\boldsymbol{\delta}})$ and $(\dot{\boldsymbol{\beta}}, \dot{\boldsymbol{\delta}})$, respectively. Then the autocorrelations and deterministic components are removed from both auxiliary estimates of $\Delta y_{it}$ and $y_{it-1}$, allowing for $\Delta \overline{y}_{it}$ and $\overline{y}_{it-1}$ to be defined as

$$\Delta \overline{y}_{it} = \Delta y_{it} + \sum_{j=1}^{p_i} \hat{\beta}_{ij} \Delta y_{it-j} + \boldsymbol{X'_{it}} \hat{\boldsymbol{\delta}}$$

$$\overline{y}_{it-1} = y_{it-1} + \sum_{j=1}^{p_i} \dot{\beta}_{ij} \Delta y_{it-j} + \boldsymbol{X'_{it}} \dot{\boldsymbol{\delta}}$$

Next, the proxies are obtained by standardizing both $\Delta \overline{y}_{it}$ and $\overline{y}_{it-1}$ by the (relevant) regression standard error, $s_i$, as follows

$$\Delta \tilde{y}_{it} = \Delta \overline{y}_{it} / s_i$$
$$\tilde{y}_{it-1} = \overline{y}_{it-1} / s_i$$

Lastly, an estimate of the coefficient $\alpha$ can be obtained from the pooled proxy equation

$$\Delta \tilde{y}_{it} = \alpha \tilde{y}_{it-1} + \eta_{it}$$

Levin, Lin, and Chu (2002) show that under the null hypothesis, a modified $t$-statistic for the resulting $\hat{\alpha}$ is asymptotically normally distributed

$$t^*_\alpha = \frac{t_\alpha - N\tilde{T}\hat{S}_N \hat{\sigma}^{-2} \mathrm{se}(\hat{\alpha}) \mu^*_{m\tilde{T}}}{\sigma^*_{m\tilde{T}}} \sim N(0,1)$$

where $t_\alpha$ is the standard $t$-statistic for $\hat{\alpha} = 0$, $\hat{\sigma}^2$ is the variance of the estimated error term, $\eta$, se($\hat{\alpha}$) is the standard error of $\hat{\alpha}$, and $\tilde{T} = T - \sum_i^N p_i / N - 1$. The remaining terms involve complicated moment calculations – see Levin, Lin, and Chu (2002) for more details.

The Breitung (2000) method differs from the Levin, Lin, and Chu (2002) method in two ways. First, only the autoregressive portion is removed when constructing the standardized proxies. Second, the proxies are transformed and de-trended. As such, the Breitung algorithm does not require kernel computations – see Breitung (2000) for more details regarding the differences in the proxies.

The Hardi (2000) panel unit root test is akin to the Kwiatkowski, Phillips, Schmidt, and Shin (1992) stationarity test, known in the literature as KPSS test, and has the null hypothesis of no unit root in any of the series in the panel. As with the KPSS test, the Hardi test is based on the residuals from the individual OLS regressions of $y_{it}$ on a constant or constant and a trend. Such a regression can be defined as

$$y_{it} = \delta_i + \eta_i t + \varepsilon_{it}$$

In particular, a LM statistic can be constructed from the residuals, $\hat{\varepsilon}$, estimated from the individual regressions

$$LM = \frac{1}{N}\left(\sum_{i=1}^{N}\left(\sum_{t} S_i(t)^2 / T^2\right) / \bar{f}_0\right)$$

where $S_i(t)$ are the cumulative sum of the residuals and $\bar{f}_0$ is the average of the individual estimators of the residual spectrum at frequency zero. Hardi (2000) then demonstrates, with mild assumptions, that

$$Z = \frac{\sqrt{N}(LM - \varepsilon)}{\zeta} \sim N(0,1)$$

where $\xi = 1/6$ and $\xi = 1/45$, if the model includes constants, and $\xi = 1/15$ and $\xi = 1/6300$, otherwise. Such a stationarity test can be considered a viable alternative to the above unit root tests because it allows the researcher to investigate the autoregressive nature of the panel in a diverse way in comparison to the ADF methodology. As such, it helps build power to the conclusions made with respect to the panel members, individually and as a group.

In contrast to the three tests described above, the Im, Pesaran, and Shin (2003) and Maddala and Wu (1999) methodologies allow for cross-sectional heterogeneity in the value of $\rho_i$. These second generation tests are in a class of their own, because of the way they combine individual unit root tests to derive a panel specific result. In the Im, Pesaran and Shin (2003) test the null and alternative hypotheses are defined as $H_0 : \rho_i = 0$ for all $i$ and $H_A : \rho_i < 0$, $i = 1,2,\ldots, N_1$, $\rho_i = 0$, $i = N_1+1$, $N_1+2,\ldots, N$. Since the $\rho_i$'s are not restricted to be identical under the null hypothesis, the alternative hypothesis is that 'not all members of the panel contain a unit root.' Once the separate individual ADF regressions have been estimated, the average of the $t$-statistics for $\alpha_i$ is then adjusted to arrive at the desired test statistics. This can be expressed as

$$\bar{t}_{NT} = \left( \sum_{i=1}^{N} t_{iT_i}(p_i) \right) / N$$

Im, Pesaran, and Shin (2003) provide simulated critical values for $\bar{t}_{NT}$ for different numbers of cross section units and series lengths, when the lag order is always zero. In the general case where the lag order is non-zero for some of the cross-sectional units, Im, Pesaran, and Shin (2003) illustrate that a properly standardized $\bar{t}_{NT}$ has an asymptotic standard normal distribution

$$W_{\bar{t}_{NT}} = \frac{\sqrt{N}\left( \bar{t}_{NT} - N^{-1}\sum_{i=1}^{N} E(t_{iT_i}(p_i)) \right)}{\sqrt{N^{-1}\sum_{i=1}^{N} Var(t_{iT_i}(p_i))}} \sim N(0,1)$$

where $E(t_{iT_i}(p_i))$ is the expected mean and Var $(t_{iT_i}(p_i)$ is the variance of the ADF regression $t$-statistics.

Maddala and Wu (1999) propose an alternative approach to panel unit root tests. In particular, they propose using Fisher's (1932) results to derive tests which combine the $p$-values from individual unit root tests. They illustrate that if $\pi_i$ is defined as the $p$-value from any individual unit root test for the cross sectional unit $i$, then under the null hypothesis of unit root for all $N$ units, an asymptotic result can be derived in the form of

$$-2\sum_{i=1}^{N} \log(\pi_i) \sim \chi^2_{2,N}$$

Additionally, Choi (2001) establishes that

$$Z = \frac{1}{\sqrt{N}} \sum_{i=1}^{N} \Theta^{-1}(\pi_i)\, \mathrm{Z} \sim N(0,1)$$

where $\Theta^{-1}$ is the inverse of the standard normal cumulative distribution function. When the Fisher tests are based on ADF regressions (refered to as 'Fisher ADF'), the number of lags used in each cross-section ADF regression must be specified. For the Phillips-

Perron (PP) form of the test (refered to as 'Fisher PP'), a kernel for estimating the frequency zero spectrum, $f_0$, must be specified by the researcher.

However, some caveats must be noted with all five unit root tests mentioned above. Breuer, McNown and Wallace (2002) point out that the alternative hypothesis in the first generation tests is rather restrictive in the sense that with as few as one stationary member in the panel, the rejection rate rises above the nominal size of the test, and increases with the number of stationary series in the panel. In such a case, the null could be correctly rejected, but the alternative of 'no unit roots' is also false in mixed panels. In contrast, the second generation tests admit that there may be a mixture of stationarity and nonstationarity contained within the panel under the alternative. However, rejection of the null in these second generation tests does not provide the researcher with information regarding the exact mix of series in the panel.

### *4.2 Panel Cointegration Tests*

If unit roots are verified in multiple variables that theoretically have a long-run relationship, then cointegration can be explored. In particular, the Pedroni (1999, 2000, 2001, 2004) cointegration methodology proposes procedures which can accommodate for considerable heterogeneity across individual members of the panel. The advantage of this approach is that it allows one to pool the long run information contained in the panel, while permitting the short run dynamics and fixed effects to be heterogeneous among different members of the panel.

In general, the following regression equation can be drawn upon to help summarize a panel cointegration test

$$y_{it} = \alpha_i + \delta_i t + \boldsymbol{X'_{it}}\, \boldsymbol{\beta_i} + e_{it}$$

where $y_{it}$ and $\boldsymbol{X_{it}}$ are both a time series panel of observable variables, with $\boldsymbol{X_{it}}$ and $\boldsymbol{\beta_i}$ being $m$-dimensional vector for each $i$. The variables are assumed to be integrated of order one for each member of the panel. Inherently, testing for cointegration amounts to testing for a unit root in the panel residuals.

The null hypothesis can be defined as $H_0$: 'all of the individuals of the panel are not cointegrated'. If the null cannot be rejected, then $e_{it}$ is also I(1). With regards to the alternative hypothesis, the researcher must first make an assumption about the underlying data generating process. If the underlying data generating process is assumed to require that all individuals of the panel be either uniformly cointegrated or uniformly not cointegrated, then the alternative hypothesis can be expressed as $H_A$: 'all of the individuals are cointegrated'. This would mean that $e$ is I(0) for all panel members. In contrast, if the underlying data generating process is assumed to permit individual members of the panel to differ in whether or not they are cointegrated, then the alternative hypothesis can be expressed as $H_A$: 'a significant portion of the individuals are cointegrated'. This can be interpreted as most of the $e_{it}$ are I(0). This follows from the parameters $\alpha_i$ and $\delta_i$ with $\beta_i$ being permitted to vary across members, which allows for the cointegrating vectors to possibly be heterogeneous across panel members.

In particular, Pedroni (1999, 2004) constructs two classes of cointegration tests. The first class is composed of four tests based on pooling the data across the within dimension of the panel. The 'panel-rho' statistic is comparable to the semiparametric 'rho' statistic studied in Phillips and Perron (1988) and Phillips and Ouliaris (1990) for the conventional time series application. Similarly, the 'panel-*t*' and 'panel-*v*' statistics are also akin to the semiparametric *t*-statistic and long run variance ratio statistic, each of which has been investigated by Phillips and Ouliaris (1990). The 'panel-ADF' statistic is constructed in a familiar fashion as the Levin, Lin, and Chu (2002) and Im, Pesaran, and Shin (2003) panel unit root test statistics, described earlier. In contrast, the second class of statistics are constructed by pooling the data along the between dimension of the panel. Therefore, these statistics in effect compute the group mean of the individual conventional time series statistics. Pedroni presents three statistics within this class, the 'group-rho,' 'group-*t*,' and 'group ADF' statistics.

All test statistics within both classes are asymptotically normally distributed. The use of these statistics is the same as for the single series case. Large positive values of the panel-*v* statistic indicate rejection of the null, whereas large negative values of the panel-rho, panel-*t*, and panel-ADF statistics indicate rejections; the same can be said for the 'group' statistics. We urge the reader to refer to Pedroni (1999, 2004), where the construction and asymptotics of the tests are outlined.

Pedroni (2000) also proposes FMOLS methods for estimating and testing hypotheses for cointegrating vectors in dynamic time series panels. Pedroni argues that the advantage of this estimator lies within its small sample properties of producing asymptotically unbiased estimators and nuisance parameter free standard normal distributions — see Pedroni (1999, p. 94) for an indepth discussion. The case is also made that, through FMOLS, inferences can be made regarding common long-run relationships, which are asymptotically invariant to the degree of short-run heterogeneity in the dynamics typically associated with panels composed of aggregate national data. Pedroni (2000) then proceeds by thoroughly outlining the underlying algorithm used to test hypotheses about common cointegrating vectors. He then also demonstrates, through monte carlo simulations, that FMOLS estimation in heterogeneous cointegrated panels has superior small sample properties and is asymptotically powerful and superconsistent — the monte carlo simulation results are found in Pedroni (2000, p. 107-114). Pedroni (2001) also points out that another advantage of this approach is that the point estimates have more useful interpretation in the event that the true cointegrating vectors are heterogeneous. As such, the FMOLS approach is appealing because it allows us to directly test the condition on the cointegrating vector that is required for long-run money demand propositions, such as the quantity theory demand for money or the Baumol-Tobin theory, to prevail.

### *4.3. Panel Data Evidence*

Table 4 reports the Levin, Lin, and Chu (LLC), Bretung (2000), Im, Pesaran, and Shin (IPS), and Fisher ADF, Fisher PP, and Hardi (2000) panel unit root test statistics preformed on the four variables of interest. The top panel displays the results for the variables in log levels and the bottom reports results for the variables in logarithmic first differences. The optimal lag length was taken to be that selected by the Akaike Information Criterion (AIC) plus 2, with the maximum lag length set equal to 2. Setting the maximum lag at 2, is common practice in the purchasing power parity and real GDP literatures when dealing with annual panel data — see, for example, Pedroni (2004) and Rapach (2002). For both monetary aggregates, the null hypothesis of a unit root in levels cannot in general be rejected at conventional significance levels. Regarding the Hardi statistic, we can reject the null hypothesis that there are no unit roots in any of the level series in the panel. Now, some ambiguity does arise when we perform the panel unit root tests on the interest rate.

When we investigate the integration properties of the interest rate variable, we find that both the Breitung and Fisher PP tests do not reject the null of a unit root, whereas the LLC, IPS and Fisher ADF tests do reject the null of a unit root. Alternatively, the Hardi test does reject the null hypothesis of no unit root. The interpretation of these mixed results leads us to the conclusion that not all panel members likely contain a unit root. This could be due to the fact that some countries conduct monetary policy via an interest rate rule rather than a money supply rule. Inspection of the raw data reveals that this is likely the case for Cyprus and Egypt — for both Cyprus and Egypt the interest rate series remains constant for most of the 1980-95 period, with minor changes after long periods of time. With the exception of the outliers, we conclude that the interest rate panel series can best be described as difference stationary. We also find support for a unit root across the panel in real output. Rapach (2002) also finds such evidence that real GDP levels are nonstationary within a panel data framework. From our perspective, the panel unit roots tests lend support to the Nelson and Plosser (1982) argument that most macroeconomic time series have a stochastic trend and are I(1). Furthermore, we interpret these results as evidence of the real business cycle theory of economic fluctuations.

TABLE 4 - *Raw Panel Unit Root Test Results in the Variables*

| Series | LLC $t^*_\alpha$-stat | Breitung $t$-stat | IPS $W$-stat | Fisher ADF[c]-stat | Fisher PP-stat | Hardi $Z$-stat[d] |
|---|---|---|---|---|---|---|
| | | | A. Log levels | | | |
| M1[a] | 1.347 | -0.540 | 4.459 | 63.997 | 53.054 | 15.199** |
| M2 | -1.537 | 1.792 | 2.443 | 87.732 | 98.591 | 15.844** |
| $R$ | -12.247** | -1.089 | -5.591** | 185.345** | 115.439 | 3.841** |
| $Y$ | -0.092 | -0.647 | 6.799 | 44.630 | 48.467 | 16.557** |
| | | | B. First differences of log levels | | | |
| M1 | -23.573** | -6.955** | -15.450** | 377.259** | 421.669** | 1.164 |
| M2 | -14.535** | -8.199** | -11.554** | 313.431** | 344.449** | 3.513** |
| $R$ | -15.389** | -6.528** | -12.800** | 333.678** | 342.167** | 4.256** |
| $Y$ | -17.337** | -7.200** | -11.946** | 295.082** | 272.224** | 1.812* |

*Notes*:

[a] Both aggregates are in real terms and logged. Standard errors are in parentheses, and * and ** indicate significance at the 5 and 1% levels, respectively.

[b] Automatic selection of lags based on AIC: 0 to 2 and a country specific constant is added to all tests.

[c] Probabilities for Fisher tests are computed using an asympotic $\chi^2$ distribution. All other tests assume asymptotic normality

[d] Newey-West bandwidth selection using Bartlett kernel.

Given that we have established evidence supportive of unit roots in the variables within the panel, we then proceed by testing for a cointegrating relationship between the variables of interest. The cointegration tests are conducted for each monetary aggregate as the dependent variable, along with both the opportunity cost and scale variables as explanatory regressors. In particular, we interpret such cointegration tests as an investigation of whether or not a long-run relationship between each of the real monetary aggregates and explanatory variables exists. To obtain the desired test statistics we utilized RATS 6.02 and Pedroni's PANCOINT source file, which is available from www.estima.com. We did consider and experiment with a couple of variants of the cointegration tests. In particular, we considered subtracting out common time effects and including heterogeneous member specific trends — although the cointegration test results with the heterogeneous member specific trends are not reported, they are available upon request from the authors. Neither of these options, however, affected the sensitivity of the conclusions drawn from each of the hypotheses tests.

Table 5 presents the results for the panel-stats and group-stats for both monetary aggregates. The panel-stats are listed in the upper portion and the group-stats are listed in the lower portion of the table. For the panel and group mean statistics we report results both for the raw data and for data that has been demeaned with respect to common time effects to accommodate some forms of cross-sectional dependency. The ADF and $t$-statistic indicate that we can reject the null hypothesis of no cointegration for all members of the panel. However, the panel-$v$, panel-rho, and group-rho statistics are always too small to reject the null hypothesis. Between all of the tests which we considered, we are left with mixed results which are typically found in the time series literature. Another explanation can be that even though all of the statistics are asymptotically consistent, they converge at different rates depending on the data generating process. In particular, Pedroni (2004) shows that with a fixed number of cross-section units and a with time dimension increasing, that the panel-$v$, panel-rho, and group-rho converge from below, indicating that they are somewhat undersized — see Pedroni (2004, p.609). As a result, we proceed as if most of the countries are cointegrated and there exist long-run equilibrium relationships which link narrow and broad real monetary aggregates to nominal interest rates and real output.

In the previous section in the initial traditional panel estimation, we determined that there was heterogeneity within the data set through diagnostic testing which indicated that the fixed effects model was preferred to the other models we considered. This finding, along with evidence of cointegration, allows us to test for the cointegrating vector using Pedroni's FMOLS procedure, which is designed explicitly for heterogeneous cointegrated panels. This source file, PANELFM, is also available on the Estima website. In particular, we can directly test whether the condition on the cointegrating vector that is required for either the classical quantity theory demand for money or the Baumol-Tobin transactions theory to hold. In the case for the quantity theory demand for money to hold, we require under the null hypothesis that interest and real output coefficients equal zero and unity, whereas under the Baumol-Tobin theory they should equal to -1/2 and 1/2, respectively. Our approach is similar to the approach taken by Pedroni in the purchasing power parity literature, where he experiments within a bivariate framework which links the logged bilateral U.S. nominal exchange rate and logged aggregate price ratio between the two countries.

TABLE 5 - *Panel and Group Cointegration Tests in the Money Demand Function*

| Monetary Aggregate[a] | $v$-stat | $\rho$-stat | $t$-stat | ADF-stat[b] |
|---|---|---|---|---|
| | | Standard panel statistics[c] | | |
| M1 | 0.987 | -0.607 | -5.455** | -6.070** |
| M2 | 1.045 | 0.961 | -2.447** | -4.003** |
| | | Time demeaned panel statistics[d] | | |
| M1 | -0.364 | 0.340 | -3.149** | -3.239** |
| M2 | 0.720 | 0.151 | -3.385** | -3.888** |
| | | Standard group statistics | | |
| M1 | — | 2.008 | -6.223** | -6.357** |
| M2 | — | 3.558 | -2.919** | -4.946** |
| | | Time demeaned group statistics | | |
| M1 | — | 2.772 | -3.333** | -3.386** |
| M2 | — | 3.007 | -3.390** | -4.366** |

*Notes*:

[a] Both aggregates are in real terms and logged. * and ** indicate significance at the 5 and 1% levels, respectively. All tests assume asymptotic normality. The critical values for the left hand 10%, 5% and 1% levels are -1.282, -1.645, and -2.326, respectively.

[b] The ADF tests use a maximum lag of 2.

[c] Panel statisticss are weighted by long run variances

[d] The time demeaned specification subtracts out the common time effect.

Finally, the FMOLS results are displayed in Table 6. We report only the group FMOLS estimates and $t$-statistics for each definition of real balances under the null hypotheses, $H_0: \beta_1 = 0$ (nominal interest elasticity equals zero) and $H_0: \beta_2 = 1$ (real income elasticity equals 1). In addition to the raw data, we again display results for data that has been time demeaned. The raw coefficient estimates and corresponding $t$-statistics for both aggregates are presented in the upper portion of the table. The time demeaned results are presented in the bottom portion of the table. It is to be noted that we do not report the individual tests because the theories we are testing are considered long-run propositions which should theoretically hold in the aggregate. However, the individual results did display a great deal of heterogeneity in the estimated slope coefficients, indicating heterogeneous cointegrating vectors. These results are available upon request.

TABLE 6 - *FMOLS Cointegrating Vector Tests*

| Variable | Coefficient | $t$-statistic |
|---|---|---|
| | M1 group results | |
| log $R$ | -0.09 | -13.81 |
| log $Y$ | 1.09 | -3.97 |
| | M2 group results | |
| log $R$ | -0.03 | -3.89 |
| log $Y$ | 1.45 | 19.49 |
| | M1 group results (time demeaned) | |
| log $R$ | -0.11 | -12.69 |
| log $Y$ | 1.15 | 3.30 |
| | M2 group results (time demeaned) | |
| log $R$ | -0.00 | 0.27 |
| log $Y$ | 1.26 | 6.52 |

*Notes:* Both aggregates are in real terms and logged $^{*}$ and $^{**}$ indicate significance at the 5 and 1% levels, respectively. All tests assume asymptotic normality. The critical values for the right hand 10%, 5%, and 1% levels are 1.282, -1.645, and -2.326, respectively. The critical values for the two-sided 10%, 5%, and 1% levels are 1.282, -1.645, and -2.326, respectively.

The results of the raw specification when we consider the narrow aggregate as the dependent variable, indicate that we cannot reject the null that the cointegrating vector contains unity but can reject the null that it contains zero ($\eta_R$=–0.09). Given the estimated interest elasticity, this finding is nearly supportive of the quantity theory demand for money and challenges the Baumol-Tobin theory. However, under the time demeaned specification, the estimated income elasticity coefficient is slightly larger and the null of unity is rejected, along with the null of the interest elasticity being equal to zero ($\eta_R$=–0.11, $\eta_Y$=1.15). Our interpretation of both specifications is that the group cointegrating vector is likely to be near, but not exact, to the hypothesized classical quantity theory prediction for the countries under investigation.

With regards to the broad cointegrating vector, the results of the raw specification indicate that we can reject both of our null hypotheses at conventional levels. The estimated real income elasticity of the demand for real broad money balances is also much higher ($\eta_R$=–0.03, $\eta_Y$=1.45) than either of the narrow estimates. The results of the time demeaned specification also reject the null hypothesis of the real income elasticity equaling unity. This estimated coefficient is much lower ($\eta_Y$=1.26), but nevertheless still hard to reconcile theoretically. However, we cannot reject the null hypothesis placed on the nominal interest elasticity of the demand for real narrow money balances. Surprisingly, the estimated coefficient is exactly equal to zero. Again, we conclude that the cointegrating vector is close to, but not exact, to the hypothesized classical prediction.

From the FMOLS results, we conclude that both the group real income and nominal interest rate elasticity coefficients are heterogeneous across aggregates, with the real income elasticity being more responsive in the broad money measure. As such, we interpret these results as an indication that the group cointegrating vector is heterogeneous across different definitions of money, even in the aggregate.

## 5. Conclusions

We have used panel data to investigate the long-run relationship between narrow and broad money, interest rates, and real GDP for 48 countries over the 1980-95 period. In particular, we considered and evaluated a variety of models. Our selection criteria and regression diagnostics indicated that the fixed effects specification, which allows for most of the heterogeneity within the 48 countries, is the ideal model among the others. However, we did find that there does exist some serial correlation.

Rather than ignoring the possible specification error, we applied new innovative panel unit root tests and found evidence that our money demand variables within the panel are for the most part I(1). As a result, this outcome then allowed us to apply panel cointegration tests. The results from these tests, in our opinion, showed evidence of a long-run relationship between money, interest rates, and real GDP. Furthermore, the application of direct panel cointegrating vector tests indicated that the quantity theory demand for money does come close holding. However, the cointegrating vector is heterogeneous not only for each individual country but for each monetary aggregate that we considered.

Evidently, as high-quality cross-country data becomes readily available from both developed and developing countries, the panel approach is an appealing methodology which can be applied to the money demand literature in order to resolve issues that have plagued researchers in this field of study for years.

Apostolos Serletis

*University of Calgary, Department of Economics, Calgary, Alberta, Canada*

Jason Vaccaro

*University of Calgary, Department of Economics, Calgary, Alberta, Canada*

## REFERENCES

Baltagi, B.H. and C. Kao (2000), *Nonstationary Panels, Cointegration in Panels and Dynamic Panels: A Survey*, in: B. Baltagi (Ed.), Advances in Econometrics, Vol. 15: Nonstationary Panels, Panel Cointegration, and Dynamic Panels, Amsterdam: JAI Press, 93-130.

Banerjee, A. (1999), "Panel Unit Roots and Cointegration: An Overview." *Oxford Bulletin of Economics and Statistics,* 61: 607-629.

Baumol, W.J. (1952), "The Transactions Demand for Cash: An Inventory Theoretic Approach." *Quarterly Journal of Economics,* 66: 545-556.

Breitung, J. (2000), *The Local Power of Some Unit Root Tests for Panel Data*, in: B. Baltagi (Ed.), Advances in Econometrics, Vol. 15: Nonstationary Panels, Panel Cointegration, and Dynamic Panels, Amsterdam: JAI Press, 161-178.

Breuer, J.B., R. McNown, and M. Wallace (2002), "Series Specific Unit Root Tests with Panel Data." *Oxford Bulletin of Economics and Statistics* 64: 527-546.

Choi, I. (2001), "Unit Root Tests for Panel Data" *Journal of International Money and Finance* 20: 249-272.

Fisher, R.A. (1932), *Statistical Methods for Research Workers*, 4th Edition. Edinburgh: Oliver & Boyd.

Fisher, M. and J. Seater (1993), "Long-Run Neutrality and Superneutrality in an ARIMA Framework." *American Economic Review,* 83: 402-15.

Fischer, A. (2005), "Measuring Income Elasticity for Swiss Money Demand: What Do the Cantons Say About Financial Innovation?" Working Paper 05.01, Swiss National Bank.

Fujiki, H., C. Hsiao and Y. Shen (2002), "Is There a Stable Money Demand Function under the Low Interest Rate Policy? A Panel Data Analysis." *Monetary and Economic Studies,* 20: 1-23.

Gujarati, D. (2003), *Basic Econometrics*, New York: McGraw-Hill/Irwin.

Hardi, K. (2000), "Testing for Stationarity in Heterogeneous Panel Data", *Econometrics Journal,* 3: 148-161.

Hsiao, C. (2003), *Analysis of Panel Data*, New York: Cambridge University Press.

Im, K.S., M.H. Pesaran and Y. Shin (2003), "Testing for Unit Roots in Heterogeneous Panels", *Journal of Econometrics,* 115: 53-74.

Kao, C. (1999), "Spurious Regression and Residual-Based Tests for Cointegration in Panel Data", *Journal of Econometrics,* 90: 1-44.

King, R. and M. Watson (1997), "Testing Long-Run Neutrality", Federal Reserve Bank of Richmond *Economic Quarterly,* 83: 69-101.

Kwiatkowski, D., P.C.B. Phillips, P. Schmidt and Y. Shin (1992), "Testing the Null Hypothesis of Stationarity Against the Alternative of a Unit Root", *Journal of Econometrics,* 54: 159-78.

Levin, A., Lin C.-F., and James Chu C.-S. (2002), "Unit Root Tests in Panel Data: Asymptotic and Finite-Sample Properties", *Journal of Econometrics,* 108: 1-24.

Maddala, G.S. and S. Wu (1999), "A Comparative Study of Unit Root Tests With Panel Data and a New Simple Test", *Oxford Bulletin of Economics and Statistics,* 61: 631-652.

McCoskey, S. and C. Kao (1998), "A Residual-Based Test for the Null of Cointegration in Panel Data", *Econometric Reviews,* 17: 57-84.

Nelson, C.R. and C.I. Plosser (1982), "Trends and Random Walks in Macro Economic Time Series", *Journal of Monetary Economics,* 10: 139-162.

Pedroni, P. (1999), "Critical Values for Cointegration Tests in Heterogeneous Cointegrated Panels", *Oxford Bulletin of Economics and Statistics,* Special Issue: 653-670.

Pedroni, P. (2000), *Fully Modified OLS for Heterogeneous Cointegrated Panels*, in: B. Baltagi (Ed), Advances in Econometrics, Vol. 15: Nonstationary Panels, Panel Cointegration, and Dynamic Panels, Amsterdam: JAI Press, 93-130.

Pedroni, P. (2001), "Purchasing Power Parity Tests in Cointegrated Panels", *Review of Economics and Statistics,* 83: 727-731.

Pedroni, P. (2004), "Panel Cointegration: Asymptotic and Finite Sample Properties of Pooled Time Series Tests With an Application to the PPP Hypothesis", *Econometric Theory,* 20: 597-625.

Phillips, P.C.B. and H. Moon (2000), "Nonstationary Panel Data Analysis: An Overview of Some Recent Developments", *Econometric* Reviews, 19: 263-286.

Phillips, P.C.B. and S. Ouliaris (1990), "Asymptotic Properties of Residual Based Tests for Cointegration." *Econometrica,* 58: 165-193.

Phillips, P.C.B. and P. Perron (1988), "Testing for a Unit Root in Time Series Regression", *Biometrica,* 75: 335-346.

Rapach, D.E. (2002), "Are Real GDP Levels Nonstationary: Evidence From Panel Data Tests", *Southern Economic Journal,* 68: 473-495.

Sarno, L. and M. Taylor (1998), "Real Exchange Rates under the Recent Float: Unequivocal Evidence of Mean Reversion", *Economics Letters,* 60: 131-137.

Serletis, A. (2007), *The Demand for Money: Theoretical and Empirical Approaches*, 2nd Edition. Springer Science.

Serletis, A. and Z. Koustas (1998), "International Evidence on the Neutrality of Money", *Journal of Money, Credit and Banking,* 30: 1-25.

Serletis, A. and Z. Koustas (2001), "Monetary Aggregation and the Neutrality of Money", *Economic Inquiry,* 39: 124-138.

Serletis, A. and J. Vaccaro (2006), "Cross-Country Evidence on the Demand for Money", *Economia Internazionale, International Economics,* 59(2), 199-224.

Sriram, S. (1999), "Survey of Literature on Demand for Money: Theoretical and Empirical Work with Special Reference to Error-Correction Models", IMF Working Paper No. 99-64.

Taylor, M. and L. Sarno (1998), "The Behavior of Real Exchange Rates During the Post-Bretton Woods Period", *Journal of International Economics,* 46: 281-312.

Tobin, J. (1956), "The Interest-Elasticity of Transactions Demand for Cash", *Review of Economics and Statistics*, 38: 241-247.

## ABSTRACT

This paper examines the demand for money using panel data for 48 countries over the 1980-95 time period. In our examination of the conventional money demand function, we begin by empirically exploiting traditional panel methodology and find support for heterogeneity among the countries. However, specification and diagnostic tests also indicate serial correlation in all of the estimated models. Recent state-of-the art advances in panel unit root and panel cointegration methodology allow us to proceed and further our analysis. Such procedures allow us to take advantage of desirable statistical properties and obtain consistent estimates in order to test long-run hypotheses.

JEL classification: C12; E41; E50

Keywords: Panel unit root tests; Panel cointegration tests; International comparisons of money demand.

## RIASSUNTO

### *Verifiche empiriche su panel data della domanda di moneta*

Questo lavoro esamina la domanda di moneta utilizzando panel data relativi a 48 paesi nel periodo 1980-1995. Il nostro esame della convenzionale funzione di domanda di moneta inizia con l'utilizzo della tradizionale metodologia panel e ottiene un risultato di eterogeneità tra i paesi.

Peraltro i test di specificazione e diagnostica indicano anche una correlazione seriale in tutti i modelli stimati. I recenti avanzamenti nelle metodologie *unit root* e di cointegrazione ci consentono di procedere ulteriormente nell'analisi. Tali metodologie ci permettono di trarre vantaggio dalle proprietà statistiche della funzione e di ottenere stime consistenti al fine di eseguire test per il lungo periodo.