Error-Correction Based Panel Estimates of the Demand for Money of Selected Asian Countries with the Extreme Bounds Analysis

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Abstract

This paper uses the extreme bounds analysis (EBA) of Leamer (1983 &1985) to analyze the robust determinants of the demand for money in a panel of 17 Asian countries for the period 1970 to 2009. These robust determinants are found to be unit root variables. Therefore, cointegration between these variables is tested with a recent time series panel method developed by Westerlund (2007). This method uses the error-correction formulation and has more power against the null of no cointegration. The results show that there is a well-defined long-run demand for money. Using the lagged error correction term from the estimated cointegrating equation, the short-run dynamic relationships are estimated. This paper, thus, suggests some useful guidelines to estimate other relationships with panel data.

Keywords: Demand for money; Extreme bounds analysis; Panel ECM; Structural breaks

JEL: C33; E41

1. Introduction

Empirical work on the demand for money is vast and some justification is necessary for yet another paper on this relationship.[1] Although this paper estimates a standard specification of the demand for narrow money ($M1$) with the time series panel data methods, it uses some recent developments in econometric methods and offers a few methodological guidelines to estimate other relationships with panel data methods.[2] The equation estimated with panel data methods, in a general form, is as follows.

$$y_{it} = \alpha + \beta_1 x_{1it} + \beta_2 x_{2it} \ldots + \beta_k x_{kit} + u_{it} \tag{1}$$

Here y is the dependent variable, $x_1 \ldots x_k$ are the explanatory variables, $u \sim N(0, \delta^2)$ is the error term and i and t are cross-section and time series dimensions respectively.[3] Although economic theory, which is about equilibrium relationships, provides important guidelines on the selection of the x variables and functional forms, it is silent on the dynamic adjustments in equation (1). This causes problems because data are generated by the real world, which is seldom in equilibrium. Therefore, how to use such data to estimate equilibrium relationships is a major problem. The London School of Economics (LSE) economists were concerned with this problem in the 1960 and 1970s and advocated the general to specific method (GETS). Its most ardent advocate is Professor David Hendry. GETS formulations argue that a variable changes due to two reasons. Firstly, if in the current period a variable is not in

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[1] For discussions related to the theoretical developments of the demand for money see Duca and VanHoose (2004), Serletis (2001), Laidler (1993a, 1993b, 1977, 1969), Barnett *et al*. (1992), Friedman (1975) and Goldfeld (1973). Many empirical studies have utilized the standard methods to analyze the stability of money demand for the industrialised and developing countries. Some recent works are Ball (2001) for USA, Lutkepohl and Wolters (1998) for Germany, Haug and Lucas (1996) for Canada, Hoffman *et al*. (1995) for five Industrial OECD countries, Artis *et al*. (1993) and Monticelli and Strauss-Kahn (1993) for seven EU countries, Rao and Kumar (2009) for 14 Asian countries, Bahmani-Oskooee and Rehman (2005) for seven Asian countries, Ghartey (1998) for Ghana, Arrau *et al.* (1995) for ten developing countries and Hafer and Kutan (1993 & 1994) for China.

[2] Analogous guidelines can also be used in the applied work with country specific time series data. By methodological guidelines we mean that these are subjective but have been used by many if not by the majority of the researchers.

[3] The error term and intercept may take different forms in the fixed effects models.

equilibrium, it moves towards its equilibrium value and partially closes the gap in the previous period between its actual and equilibrium values. This is measured by the so-called lagged error correction term (ECM_{t-1}). Secondly, a variable also changes due to any current and past changes in its determinants, but how many such lagged changes should be included is a problem. GETS approach takes the view that in the first instance it is necessary to include as many lagged changes as are necessary to avoid serial correlation in the residuals and make the equation consistent with the data generating process underlying the variable. Then, this general (unrestricted) dynamic specification can be reduced to a more parsimonious specification by deleting the insignificant changes in the variables with the variable deletion tests. PcGets software of Hendry and Krolzig (2001) is very useful for obtaining the parsimonious dynamic specification from a general dynamic specification. However, PcGets is useful only for models with country-specific time series and pure cross-section data. While the partial adjustment mechanism (PAM) is widely used for estimating empirical growth equations with panel data, following their use by Mankiw, Romer and Weil (1992) and Barro (1996), GETS specifications are not yet used in panel data estimates. Therefore, we shall use GETS in this paper.[4]

The GETS approach is developed essentially for time series models. However, Westerlund (2007) has extended it, albeit in a limited direction, to test for cointegration with panel data. There are two approaches for panel data estimation. Firstly, there is a conventional approach, which ignores the time series properties of the variables and uses the well-known fixed and random effects methods of estimation. Secondly, there are time series methods and some popular methods are due to Pedroni (2000 & 2004) and Mark and Sul (2003) and Breitung (2006) etc. Prior to estimation with the time series methods, it is necessary to test for the time series properties of the variables and, then, for cointegration. For this, there are several alternative residual based cointegration tests. However, recently Westerlund (2007) has developed an ECM based cointegration test, which has more power against the null of no cointegration. We shall apply both the conventional and Westerlund tests. Two further steps, which are ignored by many studies, are also necessary after conducting the cointegration tests. Firstly, it is necessary to estimate the cointegrating equation and secondly, it is also necessary to estimate a parsimonious dynamic

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[4] For an overview of Hendry's methodology, see Hendry (1995) and Gilbert (1986).

equation with the lagged ECM term. The cointegrating equations are estimated with the fully modified OLS (FMOLS) and the dynamic OLS (DOLS) methods usually with the RATS software. These two methods assume fixed effects with an option to include a deterministic trend.[5] However, it is not yet known how to estimate a GETS parsimonious dynamic equation with panel data. We shall argue, albeit on a heuristic basis, that this can be estimated with the standard fixed effects panel data method. An alternative method is to estimate both the cointegrating equation and the short run dynamic equation in a single step with the standard fixed effects method. This one-step method is more in line with the GETS approach. We shall report estimates with these alternative methods.

An additional methodological step in this paper is as follows. Since economic phenomena are the outcomes of human behaviour and reactions, they are likely to be volatile and the variance of the error term is likely to be generally high. A strategy to minimize this parameter is to select the variables in x that have the most robust effects in the sense of Leamer (1983). Other options are to select appropriate estimation methods. However, unless all the robust explanatory variables are included in the specification, more efficient econometric techniques alone may not be useful. Leamer's (1983) extreme bounds analysis (EBA) was used to select robust explanatory variables in the empirical growth equations by Levine and Renelt (1992) and Sala-I-Martin (1997a and 1997b). But this technique is not yet widely used by others to estimate growth equations and other empirical relationships. In this paper, we shall apply EBA to select the robust determinants of the demand for money.

With this introduction, the outline of this paper is as follows. Section 2 applies EBA to select robust explanatory variables for estimating the demand for money function with a panel of 17 Asian countries with data from 1970 to 2009. These variables are tested for unit roots and cointegration in Section 3. Section 4 reports estimates of the cointegrating equations with FMOLS, DOLS and the dynamic equations using the FMOLS, DOLS and with the one-step ECM based method. Section 5 tests for stability of the cointegrating equations and estimates the dynamic equations allowing for a single structural break. Section 6 concludes.

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[5] The Mark and Sul and Breitung methods are performed using GAUSS codes.

2. Extreme Bounds Analysis

Leamer's (1983 & 1985) *EBA* is adequately explained by Levine and Renelt (1992) and Sala-I-Martin (1997a & 1997b). Therefore, we shall be brief here. Essentially, *EBA* estimates regressions with all possible combinations of three explanatory variables at a time. In these estimates, one or two variables, usually included in many regressions, are retained as MUST variables in all combinations of the estimates. In this paper we use the log of GDP, which is a scale variable in many specifications as the MUST variable. Leamer, and Levine and Renelt have treated a variable as robust if its coefficient did not change sign in the estimates with all combinations of the three explanatory variables.[6] However, according to Sala-I-Martin this criterion is too stringent because a variable becomes fragile even if it changes sign only once. Therefore, he used the cumulative distribution functions (CDF) of the estimated coefficients to determine the robustness of the variable. He selected the 95% probability level as the critical value. Therefore, a variable becomes fragile only if its coefficient changes sign in more than 5% of the estimates. Table 1 gives *EBA* results for the robustness of five variables, which are frequently included in the specifications of the demand for money functions. Two of these variables are the log of GDP ($\ln y$) as a scale variable and the short-term rate of interest (r) as a proxy for cost of holding money. However, the rate of interest, which is often set by the central banks (financial repression), does not show much variation and its coefficient is likely to be insignificant in many time series estimates (for example see Bahmani-Oskooee and Economidou (2005) for Greece and Nielson *et al*. (2004) for Italy). Therefore, some investigators have used the rate of inflation ($\Delta \ln P$), tightness of credit conditions (CC), usually measured with the difference between the short-term rate of interest and the long term rates of interest, rl, ($CC = r - rl$) and the log of the exchange rate ($\ln FX$) as proxies for the cost of holding money.[7] When credit conditions are tight,

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[6] With this criterion, they found that only the investment rate is a robust explanatory variable in growth equations.

[7] For instance see, Bahmani-Oskooee and Rehman (2005) for Asian countries, Choi and Oxley (2004) for New Zealand, Juselius (1998) for Denmark, Ghartey (1998) for Jamaica, Bahmani-Oskooee and Shabsigh (1996) for Japan, Drake and Crystal (1994) and Baba *et al*. (1992) for USA, Fielding (1994) for African countries and Orden and Fisher (1993) for Australia and New Zealand.

the short-term rate increase relative to the long-term-rate and individuals hold less money. Therefore, the coefficient of CC is expected to be negative. The justification for including exchange rate is based on currency substitution. If it is expected that the domestic currency will depreciate, individuals hold more foreign currency and less domestic currency. It can also be argued that exchange rate movements are used as a proxy for the expected rate of inflation. If the domestic currency depreciates, given that it takes time for the completion of the exchange rate pass-through effects, the inflation rate will increase. Since the exchange rate is measured in this paper as foreign currency per unit of domestic currency, the sign of the coefficient of this variable is expected to be positive. It is also possible to use the principal component of the short-term rate of interest and these additional variables as a proxy for the cost of liquidity. Therefore, in our EBA tests we test for the robustness of the short-term rate of interest, long-term rate of interest, rate of inflation, log of the exchange rate, credit conditions, and a multiplicative term of credit conditions and the short term rate of interest ($CCr = CC \times r$) to allow for conditional effects. Results of the EBA test are in Table 1.

Table 1: Extreme Bounds Analysis with Data from 1970 to 2009

	Variable	LR CV	Average Estimated Coefficient	S-I-M CV
1	r	1	-0.0064	1
2	rl	1	-0.0059	1
3	CC	0	$0.0057E^{-2}$	0.51
4	$CC \times r$	0	$0.0045E^{-2}$	0.67
5	$\Delta \ln P$	1	-0.5369	1
6	$lnFX$	1	0.0183	1

Notes: LR CV is Levine and Renelt (1992) critical value. If it is equal to one, the variable is robust and when it is zero, the variable is fragile. S-I-M CV is Sala-I-Martin's (1997a & b) critical value. When it is ≥ 0.95 the variable is robust. CDF is the cumulative distribution of the estimates of the coefficients.

These results indicate that only CC and $CC \times r$ are fragile explanatory variables. The two interest rates, the rate of inflation and the exchange rate, in addition to the MUST $\ln y$, are all robust determinants of the demand for money. Therefore, it is to be expected that a principal component formed with these four variables will also be a robust proxy for the cost of holding money. On the basis of these EBA tests, we may specify the demand for money function as follows.

$$\ln m_{it} = \alpha + \beta_1 \ln y_{it} + \beta_2 r_{it} + \beta_3 rl_{it} + \beta_4 \Delta \ln P_{it} + \beta_5 \ln FX_{it} + u_{it} \tag{2}$$

Here $m =$ real narrow money, $y =$ real GDP, $r = 90$ day bill rate, $rl = 5$-year bond rate, $\Delta \ln P =$ rate of inflation with GDP deflator and $FX =$ exchange rate measured as US$ per unit of domestic currency. The definitions of the variables and sources of data are in the appendix.[8] It is to be expected that β_1 and $\beta_5 \geq 0$ and β_2, β_3 and $\beta_4 \leq 0.$ Other variants of equation (2) with only the short-term rate of interest and the first principal component of the four cost variables (PC) are as follows.

$$\ln m_{it} = \pi_0 + \pi_1 \ln y_{it} + \pi_2 r_{it} + \varepsilon_{it} \tag{3}$$

and

$$\ln m_{it} = \varpi_0 + \varpi_1 \ln y_{it} + \varpi_2 PC_{it} + \zeta_{it} \tag{4}$$

The error terms, u, ε, ζ of (2) to (4) are assumed to be normally distributed with zero means and constant variances. An additional specification, where ϖ_1 in equation (4) is constrained to unity, is also useful giving:

[8] Our panel data consists of 17 Asian countries $(N = 1....17)$ for the period 1970 to 2009 $(T = 1.....40)$. The selected countries are Bangladesh , Myanmar, India, Indonesia, Iran, Israel, Jordon, Korea, Kuwait, Malaysia, Nepal, Pakistan, Philippines, Singapore, Sri Lanka, Syria and Thailand. The standard panel unit root test results are provided in Table 1A in the Appendix.

$$\ln m_{it} = \mu_0 + \ln y_{it} + \mu_1 PC_{it} + \upsilon_{it}$$
$$\therefore \ln\left(\frac{m_{it}}{y_{it}}\right) = \mu_0 + \mu_1 PC_{it} + \upsilon_{it} \tag{5}$$
$$\upsilon \sim N(0, \delta_{\upsilon}^2)$$

Equation (5) is useful to test the hypothesis that money holders, in general, hold a constant proportion of real balances to income, and this may vary to changes in the cost of holding money. This specification was often used by Hendry to illustrate his GETS approach; see Hendry (1980). These additional specifications in (3) to (5) will be useful, as alternative specifications, if there is no cointegration between the variables of equation (2) due to multicolinearity.

3. Cointegration Tests

The general form of the equation, for the ECM based Westerlund (2007) cointegration tests, is as follows. For convenience, we assume only one explanatory variable and one lead and lag in its change.

$$\begin{aligned}\Delta y_{it} = &-\lambda\left[y_{it-1} - (\alpha_0 + \alpha_1 T + \beta x_{it-1})\right] \\ &+ \gamma_1 \Delta x_{it-1} + \gamma_2 \Delta x_{it+1} + \gamma_3 \Delta y_{it-1} + u_{it}\end{aligned} \tag{6}$$

The cointegration test is on whether λ is negative and significant so that the ECM (shown in the square brackets) works. The negativity of this parameter ensures negative feedback mechanism to work to drive the dependent variable towards its long run equilibrium. Two optional deterministic terms, α (intercept) and T (trend), are included in ECM for generality. One period lagged and lead changes in x are included to capture additional changes in y not due to a change in the ECM term. In reality, additional lags and leads in x may be necessary to capture more accurately the data generating process underlying the observed dependent variable. To test if the null hypothesis of no cointegration can be rejected, i.e., λ is negative and significant, Westerlund (2007) developed two group-mean tests viz., G_{τ} and G_{α}, and two analogous panel results tests, P_{τ} and P_{α}. These four test statistics are normally distributed. The first of these two tests (G_{τ}, P_{τ}) are computed with the standard errors of α, estimated in a standard way and in the second ones (G_{α}, P_{α}) are based on the Newey and West (1994) adjusted standard errors for heteroscedasticity. To overcome possible finite sample bias, bootstrap values of these four test statistics can also be generated and used in this paper. In the two group-mean based tests, the alternative hypothesis is that there is cointegration at least in one cross section unit, which is the same in many traditional panel cointegration tests. Therefore, the adjustment coefficient λ may be heterogeneous across the cross-section units. On the other hand, in the two panel data based tests, the alternative hypothesis is that adjustment to equilibrium is homogenous across cross-section units. Cointegration test results for equations (2) to (5) are in Table 2. In several estimates of these four test statistics with a deterministic trend, the null hypothesis of no cointegration could not be rejected. To conserve space these results are not reported. Table 2 report only results with a deterministic intercept.[9]

It can be seen from these results that all the four tests fail to reject the null of no cointegration for equation (2) with all the additional variables to proxy the cost of holding money. In the alternative specification (3), the null is rejected at the 10% and 5% levels by the two group-mean tests G_{τ} and G_{α}, respectively, but the two panel tests P_{τ} and P_{α} could not reject the null.[10] This implies that the adjustment coefficient λ is heterogeneous across the countries. This is confirmed by the specification in equation (4) with the first principal component for the cost variables. This implies that individual money holders use a weighted average of the short and long-term rates, rate of inflation and the exchange rate as a proxy for cost of holding money. This specification, thus seems to yield a better cointegration result because the p-values for the two group-mean tests are the lower than for equation (3). However, the specification in (5), in which the dependent variable is the inverse of the velocity of money, gave even better cointegration test results, where all the four test statistics are significant and reject the null at the 5% level. Therefore, this is our preferred equation and we shall use it for further analysis.

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[9] One period lead and lag values are used in estimation due to limited sample size. The length for the Bartlett kernel window is set at 3, which is closer to $T^{(1/3)}$. In the computation of bootstrap stand errors, 500 replications are used.

[10] Equation (3) is also tested for cointegration with the Pedroni method, see Table 2A in the Appendix.

Table 2: Westerlund Cointegration Tests 1970-2009				
Statistics	Value	Z-value	P-value	Robust P-value
$\ln m_{it} = \alpha + \beta_1 \ln y_{it} + \beta_2 r_{it} + \beta_3 rl_{it} + \beta_4 \Delta \ln P_{it} + \beta_5 \ln FX_{it} + u_{it}$ …(2)				
G_τ	-2.273	1.615	0.947	0.390
G_α	-5.120	4.929	1.000	0.948
P_τ	-7.984	1.665	0.952	0.394
P_α	-4.007	3.616	1.000	0.872
$\ln m_{it} = \pi_0 + \pi_1 \ln y_{it} + \pi_2 r_{it} + \varepsilon_{it}$ (3)				
G_τ	-2.281	-1.100	0.136	0.102*
G_α	-9.383	-0.169	0.433	0.042**
P_τ	-7.013	0.049	0.520	0.364
P_α	-6.116	-0.185	0.426	0.220
$\ln m_{it} = \varpi_0 + \varpi_1 \ln y_{it} + \varpi_2 PC_{it} + \zeta_{it}$ (4)				
G_τ	-2.399	-1.632	0.051**	0.016**
G_α	-10.298	-0.771	0.220	0.014**
P_τ	-7.253	-0.184	0.427	0.308
P_α	-6.748	-0.651	0.258	0.132
$\ln\left(\frac{m_{it}}{y_{it}}\right) = \mu_0 + \mu_1 PC_{it} + \upsilon_{it}$ (5)				
G_τ	-2.178	-1.838	0.033	0.022**
G_α	-8.488	-1.019	0.154	0.032**
P_τ	-8.834	-2.882	0.002	0.026**
P_α	-7.266	-2.819	0.002	0.024**
Note: Significance at 5% and 10% levels are denoted by ** and *, respectively.				

4. Cointegrating and Dynamic Adjustment Equations

We have estimated the coefficients of the cointegrating equations implied by equations (2) to (5) with FMOLS and DOLS.[11] They gave similar results and to conserve space we report in Table 3 the estimates of the cointegration coefficients for only equations (4) and (5). Next, as noted before, we also estimate with the one-step method both the cointegration equation and the dynamics of adjustment for (5) using the GETS approach. In line with the test results in Table 2, only a deterministic intercept term is retained in these estimates. We will discuss the GETS based one-step estimators shortly.

Table 3: Estimates of the Cointegration Coefficients 1970-2009		
$\ln m_{it} = \varpi_1 \ln y_{it} + \varpi_2 PC_{it} + \zeta_{it}$ (4)		
	Pedroni-FMOLS	Pedroni-DOLS
$\ln y$	0.930 (70.33)***	0.941 (93.65)***
PC	-0.031 (-4.93)***	-0.027 (-6.79)***
$\ln\left(\frac{m_{it}}{y_{it}}\right) = \mu_1 PC_{it} + \upsilon_{it}$ (5)		
PC	-0.034 (-5.30)***	-0.044 (-3.68)***
Notes: The reported Pedroni FMOLS and DOLS estimates are without trend. The estimates of the fixed effects intercept terms are too long and not reported. The t-ratios are reported in the parenthesis and the significance at 1% level is denoted by ***.		

[11] For an overview of the Pedroni method see Murthy (2007). Recently, Kumar and Rao (2011) have applied this method to test the validity of the Feldstein-Horioka puzzle for OECD countries.

All the estimated coefficients are significant at the 1% level and have the expected signs. Estimates of income elasticity and the semi-elasticity of cost of holding money, with FMOLS and DOLS, are close in equation (4). Income elasticity at 0.93 (FMOLS) and 0.94 (DOLS) are close to one. The coefficients of PC have the expected negative sign. In the constrained equation (5), the FMOLS estimate of the coefficient of PC is close to its estimate in equation (4). However, the DOLS estimate at -0.044 is somewhat higher in absolute value. By computing the ECM terms with these coefficients, we shall estimate the short-run dynamic equations with the classical fixed effects method. The justification is as follows. Consider the following simple specification of the short-run dynamic adjustment equation.

$$\Delta \ln m_{it} = \lambda ECM_{it-1} + \gamma_1 \Delta PC_{it} + \gamma_2 \Delta PC_{it-1} + \gamma_3 \Delta \ln m_{it-1} + u_{it} \tag{7}$$

The ECM, consisting of the levels of the variables is $I(0)$ because these variables are cointegrated. The changes in the variables are $I(0)$. Therefore, (7) can be estimated with the classical methods. To conserve space we shall use the DOLS estimates of the coefficients from Table 3 to compute ECMs.

The one-step estimator for (5) needs an explanation. The GETS specification for (5) is as follows.

$$\begin{aligned} \ln V_{it} = &-\lambda\left[\ln V_{it-1} - (\alpha + \beta PC_{1,it-1})\right] + \gamma_1 \Delta PC_{1,it-1} \\ &+ \gamma_2 \Delta PC_{1,it-1} + \gamma_3 \Delta PC_{1,it-2} + \gamma_4 \Delta \ln V_{it-1} + \eta_{it} \\ &\eta \sim N(0, \delta_\eta^2) \end{aligned} \tag{8}$$

Here V (inverse of velocity) is defined as $V = (m / y)$. The difference between (8) and (7) is that (7) uses the ECMs using the already estimated coefficients of the cointegrated variables from Table 3. In contrast, (8) estimates these coefficients and the dynamic counterpart in one-step. Like the ECM, the expression of the level variables in the square brackets is $I(0)$ because these variables are cointegrated. Therefore, (8) can be estimated with the classical panel data methods of fixed and random effects and the results are reported in Table 4.

Table 4: Estimates of the GETS Specification 1972-2009

$$\ln V_{it} = \lambda\left[\ln V_{it-1} - (\alpha + \beta PC_{1,it-1})\right] + \gamma_1 \Delta PC_{1,it} + \gamma_2 \Delta PC_{1,it-1} + \gamma_3 \Delta PC_{1,it-2} + \gamma_4 \Delta \ln V_{it-1} + u_{it}$$

	(1) FE	(2) RE	(3) FE	(4) RE	(5) FE-ECM	(6) FE-IV	(7) FE-ECM-IV
α	(see notes)	0.0033 (0.40)	(see notes)	0.0048 (0.88)	(see notes)	(see notes)	(see notes)
ECM_{t-1}	--	--	--	--	-0.1494 (6.19)	--	-0.1810 (-5.46)
λ	-0. 1756 (-4.38)***	-0.0006 (0.40)	-0.1354 (-5.39)***	-0.0007 (-0.59)	-0.1494# (6.19)	-0.1762 (5.45)	-0.1810# (-5.46)
$\lambda\beta$	-0.0119 (2.32)**	-0.0061 (2.13)**	-0.0125 (2.66)***	-0.0074 (-2.97)***		-0.0103 (1.86)*	
β $(implied)$	-0.0678	--	-0.0923	--	-0.044##	-0.0585	-0.044##
γ_1	-0.0244 (2.93)***	-0.0250 (5.75)	-0.0235 (3.17)***	-0.0240 (5.83)***	-0.0203 (3.33)	-0.0459 (1.71)*	-0.0408 (1.85)*
γ_2	-0.0025 (0.48)	-0.0072 (-1.59)	--	--	--	--	--
γ_3	0.0004 (0.11)	-0.0230 (0.53)	--	--	--	--	--
γ_4	0.0840 (1.06)	0.0140 (0.24)	--	--	--	--	--
Hausman Test RE vs FE	--	$\chi_4^2 = 134.7$ $p = 0.00$	--	$\chi_3^2 = 38.2$ $p = 0.00$	$\chi_2^2 = 46.870$ $p = 0.00$	$\chi_3^2 = 69.136$ $p = 0.00$	$\chi_2^2 = 79.633$ $p = 0.00$
LLH	749.226	--	795.763	--	793.495	736.815	736.509

Notes: The t-ratios are reported in parenthesis and the significance at 1% and 5% levels are denoted by *** and **, respectively.

In estimating (7) and (8) we started with additional lagged changes up to four periods but their coefficients were always insignificant. To conserve space estimates of (7) and (8) with all these long lags are not shown. Even the coefficients of

$\Delta PC_{1,it-1}$, $\Delta PC_{1,it-2}$ and $\Delta \ln V_{it-1}$ are insignificant, but they are reported for equation (8) in columns (1) and (2) of Table 4 for comparison. Column (1) has the fixed effects estimate and the random effects estimate is in column (2). In both estimates the coefficients of $\Delta PC_{1,it-1}$, $\Delta PC_{1,it-2}$ and $\Delta \ln V_{it-1}$ are insignificant. The ECM term in the square brackets has the correct negative sign but significant at the 5% level only in the fixed effects estimate. The Hausman test shows that the random effects estimator can be rejected in favour of the fixed effects estimator. Estimates of λ in column (1) imply that about 18% of the adjustment towards the equilibrium value of V takes place in one year. The partial elasticity coefficient with respect to the cost of holding money is -0.07.[12]

Since the coefficients of the three lagged changes are insignificant, we re-estimated this equation without these insignificant variables in columns (3) and (4) with the fixed and random effects respectively. The results show only minor changes. The adjustment coefficient in column (3) has decreased in absolute value from about -0.18 to -0.14 and the absolute value of β increased from -0.07 to -0.09. However, the log-likelihood (LLH) of the fixed effects estimator has significantly increased, implying that this parsimonious version is to be preferred.

In column (5) the two-step fixed effects estimates of equation (7), using the DOLS estimate of the coefficient for PC (-0.044 see Table 3) are shown. These are useful for comparisons between this two-step with the one-step GETS estimator in column (3).[13] In the overall comparison, the LLHs of these two estimates are very close, although LLH for the one-step estimate is slightly higher. Therefore, the GETS based one-step simpler estimator is as good as the two-step estimator. Although the estimates of the adjustment coefficient are close (-0.14 and -0.15), estimate of β is higher (-0.09) in the one-step estimate compared to the two-step estimate (-0.044).

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[12] Note that while the estimate of the joint coefficient $\lambda\beta$ and λ are significant, it is difficult to test for the significance of $\beta = (\lambda\beta / \lambda)$ because the variance for this coefficient may not exist. Therefore, we did not estimate the standard error for β.

[13] Additional lagged changes in this estimator were insignificant. Use of FMOLS estimate for the coefficient of PC did not make any significant difference. These are not shown to conserve space.

However, as shown in column (6), the higher estimate in the one-step procedure may be due to some endogeneity of ΔPC_{it}. To minimize the endogeneity of this variable, the one and two-step equations are estimated in columns (6) and (7), respectively, with the instrumental variables for ΔPC_{it}.[14] The LLHs of these estimates are close again. Notable changes are the closeness in the estimates of the adjustment coefficient λ of about -0.18 and the semi-elasticity β of about -0.06 and -0.04. The latter estimate is the DOLS estimate of the cointegrating equation in Table 3, which does not use the parsimonious specification for dynamic adjustments. Therefore, we are more inclined to claim that the GETS based one-step estimator in column (6) should be preferred for this particular sample because a different sample may need a different dynamic lag structure and may perform differently. The estimated adjustment coefficient of about -0.18 implies that it takes about 15 years for V to reach 95% of its equilibrium value.[15] At the group mean value of 0.035 for PC, a 1% increase in the cost of holding money leads to 1.7% decrease in the ratio of money balances to income. This seems a plausible estimate.

5. Breaks in the Cointegrating Equations

Some investigators have used stability tests of Hansen (1992) and Westerlund (2006) to test for the stability of estimates with panel data of nonstationary variables. However, these tests actually analyze stability of the estimates for individual cross-section units. Since there does not seem to be a test for stability of relationships estimated with all the panels, we have applied the Westerlund test for one break in the estimates of the cointegrating equation and the results are tabulated in Table 5.[16] As can be seen from these results, the relationships for the individual countries show breaks at different dates. Since structural breaks may change estimates of the panel

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[14] The instruments are ECM_{t-1}, ΔPC_{t-1}, $\Delta \ln m_{t-2}$ and $\Delta \ln y_{t-2}$.

[15] This is computed as $T_{0.95} = \left(\log\left(1 - 0.95\right) / \log\left(1 - |\lambda|\right)\right)$, where T is the number of periods necessary to complete 95% of adjustment towards equilibrium.

[16] Our choice is partly motivated due to the use of this test in Kumar and Rao (2011).

Table 5. Westerlund Tests for Structural Breaks 1970-2009 $\ln V_{it} = \alpha + \beta PC_{it}$									
Country	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Break Date	1975	2003	1985	1975	1994	1978	1975	1980	1976
Country	(10)	(11)	(12)	(13)	(14)	(15)	(16)	(17)	
Break Date	1991	1976	2003	1995	1982	1996	1977	1975	

Notes: Countries are numbered from (1) to (17) for example (1) Bangladesh, (2) Myanmar, (3) India, (4) Indonesia, (5) Iran, (6) Israel, (7) Jordon, (8) South Korea, (9) Kuwait, (10) Malaysia, (11) Nepal, (12) Pakistan, (13) Philippines, (14) Singapore, (15) Sri Lanka, (16) Syria and (17) Thailand.

cointegrating coefficients and it is cumbersome to include 17 dummies, we grouped these break dates into three broad categories viz., early (1981), mid (1985) and late (1993) breaks. It can be seen from Table 5 that in ten of these countries a break has occurred prior to 1981. However, it was not possible to select an earlier date such as 1975 or 1976 because there will be too few observations in the pre-break sample and it is not possible to estimate the cointegrating coefficients with an inadequate number of observations. Although the mid and late break dates are applicable to a small number of countries, we estimated the cointegrating coefficients for these break dates also. The estimated cointegrating coefficients are in Table 6 and they all indicate that the ratio of money holdings to GDP has increased in the post-break samples. This is to be expected due to increased monetization of transactions in the developing countries. For the early break group, DOLS estimates could not be made for the period 1970-1980 because of too few observations. The FMOLS estimates indicate that the increase in the ratio of money holdings, after the break, is 26% (shown in the parentheses in the last column). In the mid-break samples, DOLS estimates indicate a similar increase in this ratio of near 20%. In the late-break samples DOLS estimate implies that this increase is high at more than 50%. To conserve space, we have estimated the two and one-step dynamic equations only by using the FMOLS estimates for the early-break samples. This is a reasonable procedure because in the majority of these countries the break has taken place before 1981.

Table 6. Estimates of the Sub-period Cointegration Coefficients Dependent Variable: log(V)			
Break Category	Break Period	PC_1	Implied V
Early Break	FMOLS 1970-1980	-0.176 (-5.57)***	0.84
	FMOLS 1981-2009	0.087 (5.50)***	1.09 (+26%)
	DOLS 1970-1980	#	#
	DOLS 1981-2009	0.110 (5.02)***	1.12
Mid-Break	FMOLS 1970-1985	-0.114 (-7.03)***	0.89
	FMOLS 1986-2009	0.075 (3.92)***	0.93 (4%)
	DOLS 1970-1985	-0.080 (-17.62)***	0.92
	DOLS 1986-2009	0.103 (5.06)***	1.11 (19%)
Late Break	FMOLS 1970-1992	-0.127 (-5.39)***	0.89
	FMOLS 1993-2009	0.055 (3.80)***	1.05 (17%)
	DOLS 1970-1992	-0.293 (-12.04)***	0.75
	DOLS 1993-2009	0.251 (21.49)***	1.29 (54%)

Notes: The reported Pedroni FMOLS and DOLS estimates are without trend. The t-ratios are reported in the parentheses and significance at 1% level is denoted by ***. # DOLS estimates could not be made due to insufficient number of observations.

In the two-step estimator the coefficients of the two lagged ECM terms were positive and insignificant when *DPC* was instrumented as in the last two columns of Table 4. To conserve space these estimates are not reported. In our one-step estimator, the coefficients of both the lagged ECMs were negative and significant. This is reported as equation (9) below. The coefficients of additional lagged $DPCs$ and lagged dependent variable were insignificant and these are not reported.

$$\begin{aligned} DLV = &- \underset{(-4.89)^{***}}{0.1842}(1-DUM)(LV_{t-1} \; - \underset{(-1.98)^{**}}{0.0181}PC_{t-1}) \\ &- \underset{(-4.93)^{***}}{0.1829}DUM(LV_{t-1} \; - \underset{(-2.25)^{**}}{0.0193}PC_{t-1}) - \underset{(-2.19)^{**}}{0.0722}\Delta PC_t \end{aligned} \tag{9}$$

Here DUM is a dummy variable taking values of zero from 1970 to 1980 and one from 1981 to 2009. It may be noted from this estimate that the adjustment parameters did not change much. However, the long run response to the cost of holding money balances somewhat increased after 1980. This may be due to the early liberalized markets in the East Asian countries. The South and West Asian countries implemented market liberalization policies rather late in the early 1990s. From these results, it can be said that Westerlund's ECM based tests and our ECM based one-step estimates of the dynamic adjustment equation and the cointegrating equation performed far better than conventional estimates of the cointegrating equations with FMOLS and DOLS.

A final point we make is that although it is easy to perform the Granger causality tests on the individual country data, it is not clear how to perform this test in the VECM fixed effects panel data estimates. Therefore, our causality test is heuristic. If in the following equations (10) and (11), the adjustment coefficient λ is not negative and significant, then that ECM cannot explain changes in the dependent variable and the estimated cointegrating equation should not be normalized on the level of the right hand variable. On the other hand, if λ is negative and significant, then the estimated cointegrating equation can be normalized on the level of the right hand variable. This is the same as the weak exogeneity test in the country specific time series tests.

$$\Delta \ln V_t = -\lambda ECM_{t-1} + \alpha_{11}\Delta \ln V_{t-1} + \alpha_{12}\Delta PC_{t-1} \tag{10}$$

$$\Delta PC_t = -\lambda ECM_{t-1} + \alpha_{21}\Delta \ln V_{t-1} + \alpha_{22}\Delta PC_{t-1} \tag{11}$$

Estimates of the coefficients of the lagged ECMs in (10) and (11) with the t-ratios in the parentheses, respectively, are -0.1630 (-4.49) and -0.2285 (-0.48). Therefore, we may conclude that while PC weakly Granger causes V, there is no evidence to conclude that V weakly Granger causes PC, and our normalization of the estimated cointegrating equations on V is valid.

6. Conclusions

In this paper we show that the London School of Economics Hendry's General to Specific (GETS) specification can be used in panel data estimates when some assumptions are valid. We estimated with GETS the demand for money (M1) with data from a panel of 17 Asian countries over the period 1970 to 2009. Recently Westerlund (2007) extended this method to test for cointegration with panel data within the GETS approach. We have used his approach and showed how to proceed further to estimate the cointegration equations and the dynamic adjustment equations. Thus this paper extends the Westerlund framework and offers useful guidelines to estimate other relationships with panel data.

Given that a number of explanatory variables have been used in the existing studies to capture the cost of holding money, it is difficult to assert that one or another is robust. Therefore, we applied the extreme bounds analysis (EBA) of Leamer (1983 & 1985) to analyze the robust determinants of the demand for money. It is found that the nominal short-and-long-term interest rates, rate of inflation and the exchange rate, in addition to real income are all robust determinants of the demand for money; hence the principal components formed with these four variables are also robust proxies for the cost of holding money. Standard panel unit root tests revealed that these variables are I(1) in levels.

Alternative specifications of money demand are tested for cointegration with the Westerlund (2007) method. The specification in which the inverse of velocity of money depends on the first principal component gave better cointegration results where all the four test statistics, proposed by Westerlund, are significant and reject the null of no cointegration at the 5% level. The income elasticity of M1 demand is around 0.9 with the Pedroni FMOLS and DOLS methods. The coefficient of the principal component is negative and significant at the 1% level.

Further the GETS based short run dynamic equations are estimated with the classical fixed effects method. To this end, the GETS based one-step estimator (fixed effects instrumental variable) gave robust estimates. The estimated adjustment coefficient is around -0.18 and this implies that it takes about 15 years for inverse of velocity of money to reach 95% of its equilibrium value. At the group mean value of 0.035 for the principal component, a 1% increase in the cost of holding money leads to 1.7% decrease in the ratio of money balances to income in equilibrium.

Since there is no test developed to test for stability of relationships with the entire panel data, we have utilized the Westerlund (2006) structural break method. Allowing for one break in the estimates of the cointegrating equation, the relationships for the individual countries show breaks at different dates. Consequently, we grouped these break dates into three broad categories viz., early (1981), mid (1985) and late (1993) breaks and estimated for sub-sample periods with the Pedroni FMOLS and DOLS methods. The sub-sample results showed the ratio of money holdings to GDP has increased in the post-break samples. This is to be expected due to increased monetization of transactions in the developing countries. Based on our causality findings, it is inferred that while the principal component weakly Granger causes inverse of velocity of money, there is no evidence to conclude that inverse of velocity of money weakly Granger causes principal component and our normalization of the estimated cointegrating equations on inverse of velocity of money is valid.

However, there are some limitations in this paper. Our extensions are based on the empirical results, which are impressive. Nevertheless, it is necessary to validate our methods with some formal theory based approaches. We are optimistic that our extensions, which are intuitively plausible, will also hold in theory.

Data Appendix

y = Real GDP at factor cost. Data are from International Financial Statistics (IFS 2010) and World Development Indicators (2010) (WDI 2010).

r = 90 day bill-rate. Data are from IFS (2010) and WDI (2010).

m = Real narrow money supply. Data are from IFS (2010) and WDI (2010).

rl = 5-year bond rate. Data are from IFS (2010) and WDI (2010).

$\Delta \ln P =$ rate of inflation calculated from GDP deflator. Data are from IFS (2010) and WDI (2010).

$FX =$ exchange rate measured as US$ per unit of domestic currency. Data are from IFS (2010) and WDI (2010).

Appendix

Table 1A. Panel Unit Root Tests 1970-2009						
Series	LLC	Breitung	IPS	ADF	PP	Hadri
ln$m$	0.297 (0.62)	-3.268 (-0.21)	4.253 (1.00)	28.572 (0.73)	32.676 (0.53)	17.455 (0.00)***
ln$y$	1.644 (0.95)	2.412 (0.99)	1.398 (0.92)	26.556 (0.815	24.927 (0.87)	9.416 (0.00)***
r	-1.658 (0.05)**	-2.067 (0.02)**	-1.989 (0.06)*	55.125 (0.64)	43.121 (0.12)	8.483 (0.00)***
rl	-4.376 (0.37)	-1.953 (0.03)**	-3.228 (0.14)	69.609 (0.26)	25.248 (0.88)	8.058 (0.00)***
Δ ln$P$	-4.937 (0.00)***	-6.375 (0.00)***	-9.179 (0.00)***	163.122 (0.00)***	221.612 (0.00)***	5.556 (0.00)***
lnFX	-2.810 (0.02)**	1.032 (0.85)	-1.226 (0.11)	38.584 (0.27)	23.494 (0.91)	11.733 (0.00)***
PC	-3.755 (0.16)	-2.058 (0.02)**	-3.937 (0.25)	73.959 (0.04)**	89.107 (0.07)*	8.532 (0.00)***
ln$V$	0.174 (0.57)	0.758 (0.78)	-1.020 (0.15)	77.444 (0.01)***	60.512 (0.03)**	9.208 (0.00)***
Δ ln$m$	-16.002 (0.00)***	-3.290 (0.00)***	-18.076 (0.00)***	319.75 (0.00)***	1024.24 (0.00)***	4.773 (0.00)***
Δ ln$y$	-13.607 (0.00)***	-9.552 (0.00)***	-14.896 (0.00)***	246.48 (0.00)***	516.64 (0.00)***	3.730 (0.00)***
Δ r	-17.296 (0.00)***	-8.980 (0.00)***	-15.728 (0.00)***	258.94 (0.00)***	416.38 (0.00)***	3.884 (0.01)***
Δ rl	-26.665 (0.00)***	-16.407 (0.00)***	-24.205 (0.00)***	612.06 (0.00)***	155.76 (0.00)***	7.185 (0.00)***
Δ^2ln$P$	-14.963 (0.00)***	-8.457 (0.00)***	-24.671 (0.00)***	452.17 (0.00)***	2432.81 (0.00)***	3.393 (0.00)***
Δ lnFX	-13.013 (0.00)***	-9.676 (0.00)***	-13.257 (0.00)***	216.79 (0.00)***	242.38 (0.00)***	4.211 (0.00)***
Δ PC	-21.298 (0.00)***	-10.498 (0.00)***	-20.969 (0.00)***	377.66 (0.00)***	142.17 (0.00)***	3.792 (0.00)***
Δ ln$V$	-11.231 (0.00)***	-1.889 (0.00)***	-16.003 (0.00)***	282.22 (0.00)***	114.41 (0.00)***	4.738 (0.00)***

Notes: The tests are: Levin, Lin and Chu (2002, LLC), Breitung (2000), Im, Pesaran and Shin (2003, IPS), ADF Fisher χ^2 (ADF), PP Fisher χ^2 (PP), and Hadri (2001). Probability values are reported in the parentheses. ***, ** and * denotes the rejection of the null at 1%, 5% and 10% levels, respectively.

Results of the panel unit root tests are reported in Table 1A. The Hadri test in which the null is that the variable is stationary is rejected for all variables at the 5% level. The five panel unit root tests viz., LLC, Breitung, IPS, ADF and PP in which the null is that the variable is non-stationary is not rejected at 5% for ln$m$ and ln$y$. For r and rl the null is rejected by all tests at 5% level, except in the Breitung test at the 1% level. For lnFX, PC and ln$V$, the null is rejected by majority of the tests at the 5% level. Tests that reject the null at 1% level for these three variables are LLC for lnFX, Breitung and ADF for *PC* and ADF and PP for ln*V*. For Δln*P* all the tests show that it is stationary, except the Hadri test. With the exception of the Hadri test, all other tests show that the first differences of all the variables are stationary. Therefore, it is reasonable to infer that these variables are *I(1)* in levels, except Δln*P*. Since the Hadri test provides support that Δln*P* is non-stationary, we shall use Δln*P* in our cointegration analysis.

Table 2A. Panel Cointegration Tests 1970-2009

$$\ln m_{it} = \pi_0 + \pi_1 \ln y_{it} + \pi_2 r_{it} + \varepsilon_{it}$$

Test Statistic	Without trend
Panel ν - statistic	3.271***
Panel σ - statistic	-2.693**
Panel $\rho\rho$ - statistic	-3.055***
Panel ADF-statistic	-3.290***
Group σ - statistic	-1.399
Group $\rho\rho$ - statistic	-2.270**
Group ADF- statistic	-2.896**

Notes: The test statistics are distributed as *N*(0,1). *** and ** denotes significance, respectively, at 1% and 5% levels.

The conventional specification of money demand is tested for cointegration with the Pedroni method. The results are reported in Table 2A. The majority of the reported 7 tests show that there is cointegration between real M1 and its determinants (real income and nominal short-term interest rate) at the 5% level. Only the group σ test statistic is insignificant at the 5% level. Therefore, it can be concluded that these variables are cointegrated and a long run money demand function exists for the group as a whole and the members of the panel.

Table 3A: Estimates of the Cointegration Coefficients 1970-2009 $\ln m_{it} = \pi_0 + \pi_1 \ln y_{it} + \pi_2 r_{it} + \varepsilon_{it}$		
	Pedroni FMOLS	Pedroni DOLS
$\ln y$	1.013 (76.77)***	0.970 (87.65)***
r	-0.001 (-3.96)***	-0.003 (-2.61)**
Notes: The reported Pedroni FMOLS and DOLS estimates are without trend. The t-ratios are reported in parenthesis and the significance at 1% and 5% level, respectively, denoted by *** and **.		

Both methods gave consistent cointegrating estimates. The income elasticity of M1 demand is unity and statistically significant at 1% level. The semi-interest elasticity has the correct negative sign and also significant at conventional levels. The country specific income and semi-interest elasticities vary widely and this is not uncommon in the panel data studies.

References

Arrau, P., De Gregorio, J., Reinhart, C.M. and Wickham, P. (1995) 'The demand for money in developing countries: assessing the role of financial innovation,' *Journal of Development Economics*, 46, 317-340.

Artis, M. J., Bladen-Hovell, R. C. and Zhang,W. (1993) 'A European money demand function.' in: P. R. Masson and M. P. Taylor (eds.), *Policy issues in the operation of currency unions*. Cambridge: Cambridge University Press, pp. 240-263.

Baba, Y., Hendry, D.F. and Starr, R.M. (1992) 'The demand for M1 in the U.S.A., 1960–1988,' *Review of Economic Studies,* 59, 25–61.

Bahmani-Oskooee, M. and Economidou, C. (2005) 'How stable is the demand for money in Greece,' *International Economic Journal*, 19, 461-472.

Bahmani-Oskooee, M. and Rehman, H., (2005) 'Stability of the money demand function in Asian developing countries,' *Applied Economics*, 37, 773-792.

Bahmani-Oskooee, M. and Shabsigh, G. (1996) 'The demand for money in Japan: evidence from cointegration analysis, *Japan and the World Economy*, 8, 1-10.

Ball, L. (2001) 'Another look at long-run money demand,' *Journal of Monetary Economics*, 47, 31–44.

Barnett, W., Fisher, D., and Serletis, A. (1992) 'Consumer theory and the demand for money,' *Journal of Economic Literature, 30*, 2086–2119.

Barro, R. J. (1996) 'Determinants of Economic Growth: A Cross-Country Empirical Study,' NBER WP 5698, *National Bureau of Economic Research*, Cambridge, MA.

Breitung, J. (2000) 'The local power of some unit root tests for panel data,' in Baltagi, B. (ed*.), Advances in Econometrics: Non-stationary panels, panel cointegration, and dynamic panels*, JAI Press, Amsterdam, pp. 161-178.

Breitung, J. (2006) ‘A parametric approach to the estimation of cointegration vectors in panel data,’ *Econometric Reviews*, 24,151–173.

Choi, D. and Oxley, L. (2004) ‘Modelling the demand for money in New Zealand,’ *Mathematics and Computers in Simulation*, 64, 185-191.

Drake, L. and Chrystal, K.A. (1994) ‘Company-sector money demand: new evidence on the existence of a stable long-run relationship for the United Kingdom,’ *Journal of Money, Credit, and Banking,* 26, 479–494.

Duca, J.V. and VanHoose, D.D. (2004) ‘Recent developments in understanding the demand for money,’ *Journal of Economics and Business*, 56, 247-272.

Fielding, D. (1994) ‘Money demand in four African countries,’ *Journal of Economics Studies*, 21, 17-51.

Friedman, M. (1956) ‘The quantity theory of money- a restatement’ in studies in the quantity theory of money, M. Friedman, ed.,’ Chicago, IL: University of Chicago Press.

Ghartey, E.E. (1998) ‘Monetary dynamics in Ghana: evidence from cointegration, error correction modelling, and exogeneity,’ *Journal of Development Economics*, 57, 473-86.

Gilbert, C. L. (1986) ‘Professor Hendry's Econometric Methodology,’ *Oxford Bulletin of Economics and Statistics*, 48, 283-307.

Goldfeld, S. M. (1973) ‘The demand for money revisited,’ *Brookings Paper on Economic Activity*, 4, 577-638.

Hadri, K. (2000) ‘Testing for stationarity in heterogeneous panel data,’ *Econometric Journal*, 3, 148-161.

Hafer, R. W. and Kutan, A. M. (1993) ‘Price deflation, money demand, and further: evidence on money, output, and prices in China,’ *Journal of Comparative Economics,* 17, 701–9.

Hafer, R. W. and Kutan, A. M. (1994) ‘Economic reforms and long-run money demand in China: implications for monetary policy,’ *Southern Economic Journal*, 60, 936–45.

Hansen, B. E. (1992) ‘Tests and parameter instability in regressions with I(I) processes,’ *Journal of Business and Economic Statistics*, 10, 321-335.

Haug, A. A. and Lucas, R.F. (1996) ‘Long-term money demand in Canada: in search of stability,’ *The Review of Economics and Statistics,* 78, 345–48.

Hendry, D. F. (1995) ‘Dynamic Econometrics,’ Oxford: Oxford University Press.

Hendry, D.F. and Krolzig, H.M. (2001) ‘Automatic Econometric Model Selection Using PcGets,’ London: Timberlake Consultants Ltd.

Hoffman, D., Rasche, R. H. and Tieslau, M. A. (1995) ‘The stability of long-run money demand in five industrial countries,’ *Journal of Monetary Economics*, 35, 317–39.

Im, K.S., M.H. Pesaran and Y. Shin (2003) ‘Testing for unit roots in heterogeneous panels,’ *Journal of Econometrics*, 115, 53-74.

Juselius, K. (1998) ‘A structured VAR for Denmark under changing monetary regimes,’ *Journal of Business and Economic Statistics*, 16, 400-411.

Kumar, S. and Rao, B. B. (2011) ‘Time series approach to the Feldstein-Horioka puzzle with panel data from the OECD countries’, forthcoming, *The World Economy*.

Laidler, D. (1993a) ‘The Demand for Money, 4th ed.,’ New York: Harper Collins College Publishers.

Laidler, D.E.W. (1993b) 'The demand for money: theories, evidence and problems,' HarperColins College, New York.

Laidler, D. E. W. (1977) 'The demand for money: theories and evidence 2nd edition,' New York: Harper and Row.

Laidler, D. E. W. (1969) 'The demand for money: theories and evidence,' New York: Harper and Row.

Leamer, E. E. (1983) 'Let's Take the Con Out of Econometrics,' *American Economic Review* 73, 31–43.

Leamer, E. E. (1985) 'Sensitivity Analyses Would Help,' *American Economic Review* 75, 308–13.

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

Levine, R. and D. Renelt. (1992) 'A Sensitivity Analysis of Cross Country Growth Regressions,' *American Economic Review*, 82, 942-963.

Lutkepohl, H. and Wolters, J. (1998) 'A money demand system for German M3,' *Empirical Economics*, 23, 371-386.

Mankiw, N.G., D. Romer and D. Weil. (1992) 'A Contribution to the Empirics of Economic Growth,' *Quarterly Journal of Economics*, 107, 407-437.

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

Monticelli, C. and Strauss-Kahn, M-O. (1993) 'European integration and the demand for broad money,' *The Manchester School*, 61, 345-366.

Murthy, N. R. V. (2007) 'Panel cointegration analysis: An empirical example,' in Rao, B. B. (ed.) *Cointegration for the Applied Economist"*, Basingstoke, 2007: Palgrave, 2nd edition, pp.222-244.

Newey, W. K., and K. D. West. (1994) 'Automatic lag selection in covariance matrix Estimation,' *Review of Economic Studies*, 61, 631–653.

Nielsen, H., Tullio, G. and Wolters, J. (2004) 'Currency substitution and the stability of the Italian demand for money before the entry into the monetary union, 1972–1998,' *International Economics and Economic Policy*, 1, 73-85.

Orden, D. and Fisher, L.A. (1993) 'Financial deregulation and the dynamics of money, prices, and output in New Zealand and Australia,' *Journal of Money, Credit, and Banking,* 25, 273–92.

Pedroni, P. (2000) 'Fully-Modified OLS for heterogeneous cointegrated Panels,' *Advances in Econometrics,* 15, 93-130.

Pedroni, P. (2004) 'Panel cointegration: asymptotic and finite sample properties of pooled time series tests with an application to the PPP hypothesis,' *Econometric Theory,* 3, 579–625.

Rao, B.B. and Kumar, S. (2009) 'A panel data approach to the demand for money and the effects of financial reforms in the Asian countries,' *Economic Modeling*, 26, 1012-1017.

Sala-i-Martin, X. (1997a) 'I Just Ran Four Millions Regressions,' Mimeo, Columbia University.

Sala-i-Martin, X. (1997b) 'I Just Ran Two Millions Regressions,' *American Economic Review* 87, 2, 178-183.

Serletis, A. (2001) 'The demand for money- theoretical and empirical approaches,' 1st Edition, Massachusetts: Kluwer Academic Publishers.

Westerlund, J. (2007) 'Testing for error correction in panel data,' *Oxford Bulletin of Economics and Statistics*, 69, 709–748.

Westerlund, J. (2006) 'Testing for panel cointegration with multiple structural breaks,'*Oxford Bulletin of Economics and Statistics*, 68, 101-32.