Empirical Evidence for a Money Demand Function in Southeast Asian Countries: A Panel Data Analysis

Hussin Abdullah* Jachari Dahlan* and Helmi Hidthir*

ABSTRACT

Modeling demand for money is an important issue in the applied economics. The objective of this study was to estimate the demand for money in Southeast Asian countries namely Indonesia, Malaysia, Philippines, Singapore and Thailand using the panel data analysis to analysis cointegration. We constructed an aggregate data panel for Southeast Asia's five countries and used Pedroni's panel cointegration test to verify the cointegration hypothesis among the variables of the money demand function. We estimated the panel and the group mean cointegration vectors using FMOLS developed by Pedroni . We found strong evidence of cointegration among our variables.

Keywords: Demand for Money Panel Cointegration FMOLS Southeast Asian

*Economics Department, College of Arts and Sciences University Utara Malaysia

หลักฐานข้อมูลในอุปสงค์ทางการเงินของประเทศใน ภูมิภาคเอเชียตะวันออกเฉียงใต้ : การวิเคราะห์ข้อมูลจาก Panel Data

ฮัสซิน อับดุลละ* จาชาริ ดาห์ลาน* และ เฮลมิ ฮิดเตอร์*

บทคัดย่อ

รูปแบบความต้องการเกี่ยวกับการเงินเป็นประเด็นสำคัญสำหรับเศรษฐกิจในปัจจุบัน การศึกษาครั้งนี้มีวัตถุประสงค์เพื่อประมาณการความต้องการด้านการเงินในเอเชียตะวันออกเฉียงใต้ ซึ่งประกอบด้วยประเทศอินโดนีเซีย มาเลเซีย ฟิลิปปินส์ สิงคโปร์ และไทย โดยนำวิธีการวิเคราะห์ข้อมูลแบบ panel มาใช้เพื่อหาความสัมพันธ์ร่วมภายใต้ตัวแปรสถานการณ์ และเวลาที่แตกต่างกัน ซึ่งพบความสัมพันธ์ของปัจจัยดังกล่าวอย่างเด่นชัด

คำสำคัญ : อุปสงค์ทางการเงิน ความสัมพันธ์ร่วมเชิง panel FMOLS เอเชียตะวันออกเฉียงใต้

*Economics Department, College of Arts and Sciences University Utara Malaysia

Introduction

Individuals need money for their business transactions which involve goods and services. This is make transaction business for money demand. Earned income, which is unspent is for saving. Individuals must then decide how divide their savings into various kinds of assets: cash, saving accounts, government bonds, stocks, etc. This is the asset allocation or portfolio problem. Since money gains very little interest (has a low return), the higher the interest rate on non-money assets the higher the opportunity cost of holding money.

Modelling demand for money is an important issue in applied economics. The central bank's operating targeting framework has a crucial role to play in the function of demand for money. The recent direct inflation targeting framework, which is gaining more and more popularity, puts less attention on the demand for money, but still developments in the monetary aggregates belong to the set of indicators taken into account in developing countries in Southeast Asia countries.

Southeast Asian countries have carried through the development, deregulation and liberalization of their domestic money (M1 and M2) and capital markets because they have to finance the resource gap through the inflow of foreign capital or through the mobilization of domestic savings. These efforts have promoted development in the financial sectors as well as resulted in improvement in the real economy in terms of industrial structure and export components. Therefore, the debate on monetary policy in Southeast Asian is whether the money demand function has become unstable and/or the relationship between the money supply growth and inflation has weaken. If the money demand function remains stable and/or the money growth rate becomes the primary determinant of inflation, the money growth targeting becomes equivalent to inflation targeting provided the exchange rates float freely. Southeast Asian countries have experienced a declining trend in the rate of inflation and a higher degree of volatility in exchange rate among the member nations against the US dollar from 1991 to 2003.

Demand for money has received vast attention in the country specific time series studies. Developments in the unit roots and cointegration techniques and financial reforms have stimulated further empirical work on this already well-researched relationship. It is now an almost stylized fact that the demand for narrow and broad money has become temporally unstable in developed countries after continuing changes due to financial reforms. Reforms have increased competition, created additional money substitutes, increased the use of credit cards and electronic money transfers, increased liquidity of time deposits and led to higher international capital mobility (1).

Tan used quarterly data from 1973 to 1991 to estimate money demand functions, with money defined alternatively as M0, M1 and M2, to assess the role of financial developments in the structural stability of the functions. He specified the real money demands to be a function of real gross domestic product, money's own rate of return, rate of return on alternative assets, and the exchange rate index (2). Ibrahim used quarterly data covering the period 1976 to 1995 to examine the presence of long-run M1 and M2 traditional money demands. Employing residual-based cointegration tests of Engle and Granger (3), he establishes the long-run money demand for M2. However, there is no long-run relationship between the M1 demand and its determinants (4).

Rao and Kumar (1) used panel data estimation methods to estimate the cointegration for demand for money (M1) in 14 developing Asian countries. Their results show that money demand functions in these Asian countries are stable and financial reforms have yet to have any significant effects. Hosein estimated the demand for money in Iran using the autoregressive distributed lag (ARDL) approach to cointegration analysis. The empirical results show that there is a unique cointegration and stable long-run relationship among M1 monetary aggregate, income, inflation and exchange rate. He finds that income elasticity and the exchange rate coefficient are positive while inflation elasticity is negative. This indicates that depreciation of domestic currency increases the demand for money, supporting the wealth-effect argument, and people prefer to substitute physical assets for money balances that are supporting his theoretical expectation (5).

Marashdeh documents the presence of a cointegrating vector that governs the long-run money demand functions using the Johansen and Juselius procedure (6). However, his estimated dynamic models suggest the presence of structural instability in the long-run relationships between the real money demand and its determinants (7). Poole states that demand for money is stable and central banks should use money supply as the monetary policy instrument. Interest rate used as the policy instrument will only accentuate instability. Therefore, it is important to know if the money demand functions in the developing countries have become unstable. Stable money demand implies that using the rate of interest as the monetary policy instrument is inappropriate (8).

Hussain et al. tested a stable long-run money demand function for Pakistan which experienced high inflation. The money demand function includes real money balance, real GDP, the rate of inflation, interest rate on time deposits and/or the exchange rate and the financial innovation. The empirical results support the long-run stable demand function for money in Pakistan. All variables in the money demand equation are individually significant and the signs are as expected (9).

Dahalan *et al.* investigated the appropriate scale variable for money demand for Malaysia for four monetary aggregates. These are considered for simple sums M1, and M2 and for Divisia M1 and M2. They found that the non-nested tests for the M1 aggregates were inconclusive but more support is given for income as the scale variable for the M2 aggregates, especially for Divisia M2. The other procedures also show that income is preferred for M2 and some support is also given for income for M1 aggregates (10).

Our main objective of this paper is to highlight the cointegrating properties of M1 and M2 monetary aggregates, income, inflation, exchange rate, and interest rate using the cointegrating technique known as Pedroni's approach.

The rest of this paper is organized as follows: section two introduces the theoretical model and discusses Pedroni's approach, section three displays major results, and section four concludes.

The Specification of Model and the Methodology

According to Harb in the literature the money demand is formulated as follow (11):

$$M_t = f(S_t, O_t) \qquad f_s' > 0, \quad f_o' < 0 \quad [1]$$

where M_t is the quantity demanded of money, S_t is a scale variable and O_t is the opportunity cost of holding money. The variables used for the estimation of the money demand function depend on the theoretical function of money.

We estimate the following money demand equation

$$M_{i,t} = \beta_{i,0} + \beta_{i,1}Y_{i,t} + \beta_{i,2}R_{i,t} + \beta_{i,3}ER_{i,t} + \beta_{i,4}\pi_{i,t} + \varepsilon_{i,t} \quad [2]$$

where the *i* index refers to the given number of the panel. $M_{i,t}$ is the natural logarithm of the real monetary aggregate for country *i*. In our study we used M1 and M2 as monetary aggregates. $Y_{i,t}$ is the logarithm of the real income as a scale variable; $R_{i,t}$ is the local real interest rate; $ER_{i,t}$ is the real exchange rate per US dollar, a depreciation of the domestic currency or increase in ER raises the value of the foreign assets in term of domestic currency, and $\pi_{i,t}$ is the rate of inflation. β_0, $\beta_{1,\ldots,}$ β_4 are the slopes to be estimated.

The five series in the model are expected to be non-stationary. Thus, we rewrite the previous equation as follows

$$\varepsilon_{i,t} = M_{i,t} - \beta_{i,0} - \beta_{i,1}Y_{i,t} - \beta_{i,2}R_{i,t} - \beta_{i,3}ER_{i,t} - \beta_{i,4}\pi_{i,t} \quad [3]$$

Through Equation (3) we suggest that the five variables must be cointegrated with the cointegrating vector $[1, \beta_{i,1}, \beta_{i,2}, . \beta_{i,3}, \beta_{i,4}]$. Imposing a homogenous vector across the panel led to dramatic consequences. We used the hetegenous panel tests and estimations techniques as described in Pedroni (14, 16, 21). Those techniques are detailed in the following section.

In order to investigate the possibility of panel cointegration, it is first necessary to determine the existence of unit roots in the data series. There are several unit root tests specifically for panel data which have been introduced in past decades. Each panel unit root test data has its own benefits and limitations and for this study we have chosen the Levin, Lin and Chu (LLC, hereafter), Im, Pesaran and Shin (IPS, hereafter) and Maddala and Wu (MW, hereafter) tests, which are based on the well-known Dickey-Fuller procedure (12-14).

We started with LLC which found that the main hypothesis of panel unit root is as follows:

$$\Delta y_{it} = \Phi_i y_{i,t-1} + \sum_{L=1}^{p_i} \rho_{i,l} \Delta y_{i,t-L} + \varepsilon_{i,t}$$

$m = 1, 2, \ldots$ [4]

where $y_{i,t}$ refers to variable $M_{i,t}$, $Y_{i,t}$, $R_{i,t}$ $ER_{i,t}$, $\dot{\pi}_{i,t}$ and Δ refers to the first difference. The hypothesis test is $H_0 : \Phi_i = 0$ for existence of unit root whereas $H_a : \Phi_i < 0$ for all i for nonexistence of unit root. As p_i is unknown, LLC suggests a three-step procedure in the test. In the first step, we obtain the ADF regression which has been separated for each individual in the panel and generate two orthogonalized residuals. The second step requires an estimation of the ratio of long-run to short-run innovation standard deviation for each individual. The last step requires us to compute the pooled t-statistics.

Im, Pesaran and Shin, denoted as IPS proposed a test for the presence of unit roots in panels that combine information from the time series dimension with that from the cross-section dimension, such that fewer time observations are required for the test to have power. Since the IPS test has been found to have superior test power by researchers in economics to analyze long-run relationships in panel data, we will also employ this procedure in this study. IPS begin by specifying a separate ADF regression for each cross-section with individual effects and no time trend:(13)

$$\Delta y_{it} = \alpha_i + \rho_i y_{i,t-1} + \sum_{j=1}^{p_i} \beta_{ij} \Delta y_{i,t-j} + \varepsilon_{it} \quad [5]$$

where $i = 1, \ldots, N$ and $t = 1, \ldots, T$

IPS use separate unit root tests for the N cross-section units. Their test is based on the Augmented Dickey-fuller (ADF) statistics averaged across groups (12). After estimating the separate ADF regressions, the average of the t-statistics for p_1 from the individual ADF regressions, $t_{iT_i}(p_i)$ as:

$$\bar{t}_{NT} = \frac{1}{N} \sum_{i=1}^{N} t_{iT}(p_i \beta_i) \quad [6]$$

The *t*-bar is then standardized and it is shown that the standardized *t-bar* statistic converges to the standard normal distribution as N and $T \rightarrow \infty$. IPS (1997) showed that *t-bar* test has better performance when N and T are small. They proposed a cross-sectionally demeaned version of both tests to be used in the case where the errors in different regressions contain a common time-specific component.

Finally, Maddala and Wu, denoted as MW developed a test based on the probability values of all root unit individual tests. An alternative approach to panel unit root tests uses Fisher's results to derive tests that combine the *p*-values from individual unit root tests (15). The statistic is given as:

$$2\sum_{i=1}^{N}\log(\pi_i) \rightarrow \chi^2_{2N} \quad [7]$$

where π_i is the p-value of the test statistic in unit *i*, and is distributed as a $\chi^2(2N)$ under the usual assumption of cross-sectional independence. When the Fisher test is based on ADF test statistics, we must specify the number of lags used in each cross-section ADF regression. Maddala and Wu showed that it is more powerful than the *t*-bar in the IPS test.

Panel Cointegration Tests

The next step is to test for the existence of a long-run cointegration among real per capita GDP growth rates and the independent variables using panel cointegration tests suggested by Pedroni (16, 17). We will make use of seven panel cointegrations by Pedroni, since he determines the appropriateness of the tests to be applied to estimated residuals from a cointegration regression after normalizing the panel statistics with correction terms.

The procedures proposed by Pedroni make use of estimated residual from the hypothesized long-run regression of the following form (14):

$$y_{i,t} = \alpha_i + \delta_i t + \beta_{1i}x_{1i,t} + \beta_{2i}x_{2i,t} + \mathrm{K} + \beta_{Mi}x_{Mi,t} + e_{i,t} \quad [8]$$

for t = 1,.....,*T; i* = 1,....,*N; m* = 1, ..., *M,*

where *T* is the number of observations over time, *N* the number of cross-sectional units in the panel, and *M* the number of regressors. In this set up, α_i is the member specific intercept or fixed effects parameter which varies across individual cross-sectional units. The same is true of the slope coefficients and member specific time effects, $\delta_i t$.

Pedroni proposes the heterogeneous panel and the heterogeneous group mean panel test statistics to test for panel cointegration. He defines two sets of statistics. The first set of three statistics $Z_{\hat{v},N,T}$, $Z_{\hat{\rho}N,T\ 1}$ and $Z_{tN,T}$ is based on pooling the residuals along the within dimension of the panel (16, 15). The statistics are as follows:

[9]

$$Z_{\hat{\rho}N,T\ 1} = T\sqrt{N}\left(\sum_{i=1}^{N}\sum_{t=1}^{T}\hat{L}_{11i}^{\ 2}\hat{e}_{i,t\ 1}^{2}\right)^{1}\sum_{i=1}^{N}\sum_{t=1}^{T}\hat{L}_{11i}^{\ 2}\left(\hat{e}_{i,t\ 1}\Delta\hat{e}_{i,t}\quad\hat{\lambda}_{i}\right) \qquad [10]$$

$$Z_{tN,T} = \left(\tilde{\sigma}_{N,T}^{2}\sum_{I=1}^{N}\sum_{T=1}^{T}\hat{L}_{11i}^{\ 2}\hat{e}_{i,t\ 1}^{2}\right)^{1/2}\sum_{i=1}^{N}\sum_{t=1}^{T}\hat{L}_{11i}^{\ 2}\hat{e}_{i,t\ 1}^{2}\left(\hat{e}_{i,t\ 1}\Delta\hat{e}_{i,t}\quad\hat{\lambda}_{i}\right) \qquad [11]$$

where $\hat{e}_{i,t\ 1}$ is the residual vector of the OLS estimation of Equation (8) and where the other terms are properly defined in Pedroni (18).

The second set of statistics is based on pooling the residuals along the between dimension of the panel. It allows for a heterogeneous autocorrelation parameter across members. The statistics are as follows:

$$\tilde{Z}_{\hat{\rho}N,T\ 1} = \sum_{i=1}^{N}\left(\sum_{t=1}^{T}\hat{e}_{i,t\ 1}^{2}\right)^{1}\sum_{t=1}^{T}\left(\hat{e}_{i,t\ 1}\Delta\hat{e}_{i,t}\quad\hat{\lambda}_{i}\right) \qquad [12]$$

$$\tilde{Z}_{tN,T\ 1} = \sum_{i=1}^{N}\left(\sum_{t=1}^{T}\hat{e}_{i,t\ 1}^{2}\right)^{1/2}\sum_{t=1}^{T}\left(\hat{e}_{i,t\ 1}\Delta\hat{e}_{i,t}\quad\hat{\lambda}_{i}\right) \qquad [13]$$

These statistics compute the group mean of the individual conventional time series statistics. The asymptotic distribution of each of those five statistics can be expressed in the following form:

$$\frac{X_{N,T}\quad\mu\sqrt{N}}{\sqrt{v}} \Rightarrow N(0,1) \qquad [14]$$

where $X_{N,T}$ is the corresponding form of the test statistics, while μ and are the mean and variance of each test respectively. They are given in Table 2 in Pedroni (16). Under the alternative hypothesis, Panel *v* statistics diverges to positive infinity. Therefore, it is a one sided test where large positive values reject the null of no cointegration. The remaining statistics diverge to negative infinity, which means that large negative values reject the null.

Fully Modified Ordinary Least Squares (FMOLS) Estimation

In this section we adopt the FMOLS procedure from Christopoulos and Tsionas (19 and 20). In order to obtain asymptotically efficient consistent estimates in the panel series, non-exogeneity and serial correlation problems are tackled by employing fully-modified OLS (FMOLS) introduced by Pedroni (21). Since the explanatory variables are cointegrated with a time trend, and thus a long-run equilibrium relationship exists among these variables through the panel unit root test and panel cointegration test, we proceed to estimate the Equation (2) by the method of fully-modified OLS (FMOLS) for heterogenous cointegrated panels (21, 22). This methodology allows consistent and efficient estimation of cointegration vector and also addresses the problem of nonstationary regressors, as well as the problem of simultaneity biases. It is well known that OLS estimation yields biased results because the regressors are endogenously determined in the *I*(1) case. The starting point is the OLS as in the following cointegrated system for panel data:

$$y_{it} = \alpha_i + x_{it}'\beta + e_{it} \qquad [15]$$

$$x_{it} = x_{i,t\ 1} + \varepsilon_{it}$$

where $\xi_{it} = [e_{it}, \varepsilon_{it}']$ is the stationary with covariance matrix Ω_i. The estimator β will be consistent when the error process $\omega_{it} + [e_{it}, \varepsilon_{it}']'$ satisfies the assumption of cointegration between y_{it} and x_{it}. The limiting distribution of OLS estimator depends upon nuisance parameters. According to Phillips and Hansen, a semi-parametric correction can be made to the OLS estimator that eliminates the second order bias caused by the fact that the regressors are endogenous (23). Pedroni follows the same principle in the panel data context, and allows for the heterogeneity in the short-run dynamics and the fixed effects (21, 22). Pedroni's FMOLS estimator is constructed as follow:

$$\hat{\beta}_{FM} \quad \beta = \left(\sum_{i=1}^{N} \hat{\Omega}_{22i}^{\ 2} \sum_{t=1}^{T} (x_{it} \quad \hat{x}_t)^2 \right)^{1}$$

$$\sum_{i=1}^{N} \hat{\Omega}_{11i}^{\ 1} \hat{\Omega}_{22i}^{\ 1} \left(\sum_{t=1}^{T} (x_{it} \quad \bar{x}_t) e_{it} \quad T\hat{\gamma}_i \right) \qquad [16]$$

$$\hat{e}_{it} = e_{it} \quad \hat{\Omega}_{22i}^{\ 1} \hat{\Omega}_{21i}, \qquad \hat{\gamma}_i = \hat{\Gamma}_{21i} + \hat{\Omega}_{21i}^{0}$$

$$\hat{\Omega}_{22i}^{\ 1} \hat{\Omega}_{21i} \left(\hat{\Gamma}_{22i} + \hat{\Omega}_{22i}^{0} \right)$$

where the covariance matrix can be decomposed as $\Omega_i = \Omega_i^0 + \Gamma_i + \Gamma_i$ where Ω_i^0 is the contemporaneous covariance matrix, and Γ_i is a weighted sum of autocovariances. Also, $\hat{\Omega}_i^0$ denotes an appropriate estimator of Ω_i^0.

In this study, we employed the panel group FMOLS test from Pedroni (1996, 2000). An important advantage of the panel group estimators is that the form in which the data is pooled allows for greater flexibility in the presence of heterogeneity of the cointegrating vectors. Test statistics constructed from the panel group estimators are designed to test the null hypothesis $H_0 : \beta_i = \beta_0$ for all *i* against the alternative hypothesis $H_A : \beta_i \neq \beta_0$, so that the values for β_i are not constrained to be the same under the alternative hypothesis. Clearly, this is an important advantage for applications such as the present one, because there is no reason to believe that, if the cointegrating slopes are not equal to one, they necessarily take on some other arbitrary common value. Another advantage of the panel group estimators is that the point estimates have a more useful interpretation in the event that the true cointegrating vectors are heterogeneous. Specifically, point estimates for the panel group estimator can be interpreted as the mean value for the cointegrating vectors (18).

Results

The data used in the present article are from annual reports and they cover the five member countries Southeast Asia and run from 1977 to 2007. Therefore, we have 30 annual observations for each member.

The following subsection (3.1) discusses the existence of the unit root in the data series while subsection (3.2) and (3.3) discuss the major results of the paper.

Unit Root Test

As mentioned before, we used the LLC, the IPS, and the MW test to verify the existence of unit root in the panel series. The results are shown in Table 1. Our variables are not stationary in the constant without time trend level by applying the LLC, IPS and MW tests which are also applied for heterogeneous panels to test the series for the presence of a unit root. The results of the panel unit root tests confirm that the variables are non-stationary at the level.

The last three columns show that in the first order differences of all the series the null hypothesis of unit root test is rejected at 95% critical value (1% significant level). Hence, in the different tests based on the LLC, the IPS, and the MW tests, there is evidence that all the series are in fact integrated of order one. The next step would be to test for cointegration among our set of variables.

Table 1 Panel unit root test 1977-2007

Variable	Level			First order difference		
	Constant without trend			Constant + Trend		
	LLC	IPS	MW	LLC	IPS	MW
M1	-0.322	1.249	6.333	-6.498*	-7.502*	62.912*
M2	-0.068	-0.812	10.201	-7.613*	-12.579*	115.930*
GDP	0.518	-1.167	10.025	-9.992*	-12.231*	113.714*
Interest	-0.847	0.179	10.958	-5.838*	-6.242*	53.073*
Exchange Rate	1.098	-0.532	11.710	-9.271*	-8.587*	71.829*
Inflation	0.154	-1.059	9.326	-11.218*	-13.386*	124.209*

Note: LLC - *Levin, Lin and Chu*, IPS - *Im, Pesaran and Shin*, MW - *Maddala and Wu*

*, ** indicates rejection of the null hypothesis of non-cointegration at 1% and 5%, levels of significance

Cointegration Analysis

The cointegration analysis is the step to test whether the variables are cointegrated using Pedroni's procedure (15,16,18). This is to investigate whether a long-run steady state or cointegration exists among the variables and to confirm the statements of Oh *et al*. (1999) and Coiteux and Olivier (2000) that the panel cointegration tests have much higher testing power than conventional cointegration tests. Since the variables are found to be integrated in the same order *I*(1), we continued with the panel cointegration tests. Cointegration tests are carried out for constant and constant plus time trend, and the summary of the results of the cointegrations analyses are presented in Table 2.

In constant level, we found that M1 indicates that none of the seven statistics reject the null hypothesis of non-cointegration at the 1% level of significance, while in M2 the seven statistics reject the null hypothesis. In the panel cointegration test for M1 and M2 with constant plus trend level, the results indicated that six out of seven statistics reject the null hypothesis of non-cointegration at the 1% and 5% levels of significance. It is shown that independent variables do hold cointegration in the long run for a group of five Southeast Asian countries with respect to monetary aggregate. However, since all the statistics concluded in favour of cointegration, and this, combined with the fact that according to Pedroni the panel non-parametric (*t*-statistic) and parametric (*adf*-statistic) statistics are more reliable in constant plus time trend, we conclude that there is a long-run cointegration among our variables in the five Southeast Asian countries (14).

Table 2 The Pedroni Panel Cointegration Test, Monetary aggregate: M1&M2

Test	Constant without trend		Constant + Trend	
	M1	M2	M1	M2
Panel *v*-Statistic	-0.758	4.403*	-0.855	3.909*
Panel ρ-Statistic	0.593	-3.038*	-2.024*	-1.769*
Panel *t*-Statistic: (non-parametric)	-0.399	-6.025*	-1.309**	-5.279*
Panel *t*-Statistic: (parametric)	-0.268	-4.830*	-2.212*	-5.034*
Group ρ–Statistic	1.045	-2.437*	-2.673*	-1.081
Group *t*-Statistic: (non-parametric)	-0.550	-6.746*	-1.494**	-5.476*
Group *t*-Statistic: (parametric)	-0.377	-5.303*	-2.384*	-5.198*

Note: All statistics are from Pedroni's procedure (1999) where the adjusted values can be compared to the N(0,1) distribution. The Pedroni (2004) statistics are one-sided tests with a critical value of -1.64 ($k < -1.64$ implies rejection of the null), except the *v*-statistic that has a critical value of 1.64 ($k > 1.64$ suggests rejection of the null).

*, ** indicates rejection of the null hypothesis of non-cointegration at 1% and 5%, levels of significance

FMOLS Estimation

As mentioned before, since the OLS estimator yields a biased distribution of the residuals, we use FMOLS methodology proposed by Pedroni to estimate the panel cointegration vector (21). FMOLS is superior to OLS when applied to heterogeneous panel with (1) variables. Its distribution is standard and is asymptotically unbiased and free of nuisance parameters. FMOLS results in consistent standard error and therefore consistent t-statistics. FMOLS also allow for heterogeneity by allowing the associated serial correlation properties of the error processes to vary across members of the panel (10).

The results of the regressions estimations are shown in Table 3. Real M1 and M2 were used as monetary aggregate. In Indonesia and Thailand, we found that elasticity with respect to the interest rate was positive [0.40 (M1 and M2), 1.06 (M1) and 1.05 (M2), respectively] and statistically significant at the 5% level. Elasticity with respect to the exchange rate was negative

Table 3 FMOLS regression - with time dummies

	Monetary aggregate = M1				Monetary aggregate = M2			
Country	GDP	Interest	Ex. rate	Inflation rate	GDP	Interest	Ex. rate	Inflation rate
Indonesia	0.33 (0.66)	0.40* (2.67)	-0.48* (-3.02)	0.66 (1.19)	0.31 (0.66)	0.40* (2.79)	-0.47* (-3.17)	0.60 (1.15)
Malaysia	1.38* (2.83)	-0.40 (-1.60)	-3.98* (-3.93)	1.73* (3.56)	1.83* (4.22)	-0.49* (-2.18)	-4.05* (-5.62)	2.17* (5.00)
Philippines	-0.12 (-0.22)	0.09 (0.23)	-0.25 (-1.31)	0.33 (0.58)	-0.20 (-0.37)	0.11 (0.26)	-0.26 (-1.40)	0.21 (0.37)
Singapore	1.08 (1.40)	0.89* (2.93)	0.99 (1.57)	1.11 (1.41)	0.90 (1.24)	0.84* (2.92)	0.83 (1.40)	0.91 (1.23)
Thailand	0.10 (0.77)	1.06 *(8.23)	-0.84* (-8.14)	0.16 (1.06)	0.09 (0.71)	1.05* (8.57)	-0.85* (-8.62)	0.15 (1.05)
Panel Group	0.55* (2.43)	0.41* (5.57)	-0.75* (-6.63)	0.80* (3.49)	0.59* (2.89)	0.38* (5.53)	-0.96* (-7.79)	0.81* (3.93)

Note: The null hypothesis for the *t*-ratio is $H_0 = \beta_i = 0$; Figures in parentheses are *t*-statistics. (*) and (**) significant with 95% (90%) confidence level

(-0.48 and -0.47, -0.84 and 0.85, respectively) and statistically significant at the 5% level for M1 and M2. The GDP and inflation rate elasticity in case of Indonesia were not significant but positive. In Malaysia, M1 and M2 showed good performances in our research. All the independent variables were positive and were elasticity negative and statistically significant at the 5% level. But elasticity with respect to the interest rate for M1 was not significant. We also found that in the Philippines all the independent variables were statistically not significant for M1 and M2. Only the interest rate elasticity in the case of Singapore were statistically significant and positive (0.84 and 0.89, respectively) for M1 and M2.

The pool panel estimation is shown in the bottom row of Table 3. The panel estimation pools the data of all members, but allows for heterogeneous serial correlation properties across members. The pooled panel t-statistic can be used to test H_0: $\beta_i = \beta_0$ for all *i* versus H_1: $\beta_i = \beta_a \neq \beta_0$ where β_0 is the hypothesized common value for β under the null and β_a is an alternative common value (10). The pooled estimator shows that the elasticity of the variables is positive and statistically significant at the 5% level except for the exchange rate which is negative and statistically significant at the 5% level.

Discussion

In this paper we have shown that the variables in the demand for money in five Southeast Asian countries are non-stationary in their levels in the panel series. The existence of a valid long-run money demand function is still important for the conduct of monetary policy. The LLC, the IPS and the MW tests for panel unit roots support the view that all the variables appearing on a standard money demand function are $I(1)$. These stationary tests show that a long-run relationship exists between the dependent and the independent variables. We proceeded the next step with Pedroni's panel cointegration test to verify the cointegration hypothesis among the variables of the money demand function. We estimated the panel and group means cointegration vectors using FMOLS developed by Pedroni (22).

In Pedroni's panel cointegration test, monetary aggregate M2 shows a very good performance at constant and constant plus time trend. This means that the real M2 is a predictable monetary aggregate compared to the monetary aggregate M1. In order to identify a long-run stable money demand function for M1 and M2 for individual countries, we used FMOLS and the results showed that the real income affects money demand for monetary aggregate M1 and M2 only for Malaysia as an individual country. As a group, we found that the real income was sufficient for the formulation of a long-run stable demand for M1 and M2 money in the Southeast Asian countries.

The expected rate of inflation is usually the only variable used as the opportunity cost of holding money. From our research in the five Southeast Asian countries, we found that only Malaysia had a strong incentive for persons to switch out of money and into real assets when there are strong inflationary expectations. This rate of inflation is statistically significant in explaining changes in the demand for money. Other countries showed that there were positive relationships between inflation expectation and monetary aggregate M1 and M2 but were not statistically significant.

The estimated coefficient for the interest rate spread in this model was highly significant and had consistent signs for Indonesia, Malaysia, Singapore and Thailand. The estimated coefficient for exchange rate was highly significant with monetary aggregate M1 and M2 for Indonesia, Malaysia and Thailand.

Based on the results, we conclude that a broad definition of money M2 is a better measure than a narrow definition of money M1 in considering the long-run economic impacts of changes in monetary policy in Southeast Asian countries. Even though our results are fairly significant and imply a stable demand for money in the panel of the five Southeast Asian countries, they have some limitations. We do not have better methods to decide endogenous in the panel data. We hope that our paper will provide incentives for further work to improve panel data estimation methods.

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