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A panel data approach to the demand for money and the effects of financial reforms in the Asian countries

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ABSTRACT

Alternative panel data estimation methods are used to estimate the cointegrating equations for the demand for money (M1) for a panel of 14 Asian countries from 1970 to 2005. The effects of financial reforms are analyzed with estimates for two sets of sub-samples and two break dates. Our results show that money demand function has been stable and financial reforms are yet to have any significant effects. Since there is no evidence for instability in the demand for money, the central banks of these countries should use money supply, instead of the rate of interest, as the monetary policy instrument.

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1. Introduction

Demand for money and its stability have 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 the developed countries after the continuing changes to the financial sector due to financial reforms. Reforms have increased competition, created additional money substitutes, increased use of credit cards and electronic money transfers, increased liquidity of the time deposits and lead to higher international capital mobility. Consequently many central banks of the developed countries have abandoned money supply as a policy instrument because it is difficult to predict demand for money with a temporally unstable function. Furthermore, the Taylor rule has made attractive the use of the bank rate as the policy instrument by arguing that it will enhance the built-in stability of the economy. Therefore, since the late 1970s many central banks in the developed countries have abandoned using money supply as the policy instrument and switched to adjusting the rate of interest to stabilize the economy. This is also consistent with Poole (1970) who showed that the rate of interest should be targeted if demand for money is unstable.[2]

Following these developments, central banks in many developing countries have also started using the rate of interest as their monetary policy instrument although there is no convincing evidence that their money demand functions have become unstable after financial reforms. Bahmani-Oskooee and Rehman (2005) found that demand for money functions in several developing Asian countries, by and large, are stable.[3] According to Poole (1970) if demand for money is

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2 Poole's arguments are well explained in Mishkin (2003) pp.459–463. However, Rao (2007) has argued that central banks should not be given the power to change the interest rates because such changes have significant distributional effects. The recent worldwide severe downturns seem to be due to artificially lowering the interest rate by the FED in the USA. If interest rates were left to be determined by the market forces, perhaps we could have avoided credit bubbles, accumulation of huge toxic loans and severe downturns.

3 The countries selected in this study are India, Indonesia, Malaysia, Pakistan, the Philippines, Singapore and Thailand. They found that while in India, Indonesia and Singapore, demand for M1 is stable, in Malaysia, Pakistan, the Philippines and Thailand demand for broad money (M2) is stable. In the latter 4 countries the cointegrating equations with M1 are not well determined. However, in a recent paper Sumner (2009) with data from 1950 to 1998 showed that the components of the demand for money (M1 and M2) in Thailand have been stable and well determined. In our paper we shall use M1 for analysis because M1 is by and large the dominant component of money supply in the developing countries. We do not rule out using M2 or M3 as alternatives and our framework can be easily used for M2 or M3. In particular it is hard to control NBFI in the organized and unorganized sectors in the developing countries and hence our preference for the narrow definition of money.

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Table 1
Panel unit root tests 1970–2005.

Series	LLC	Breitung	IPS	ADF	PP	Hadri
ln (*M*)	−1.977 (0.02)*	2.461 (0.99)	−2.061 (0.02)*	52.132 (0.003)*	54.082 (0.00)*	7.700 (0.00)*
ln (*Y*)	1.883 (0.97)	−3.628 (0.00)*	1.256 (0.90)	24.621 (0.65)	25.440 (0.60)	5.509 (0.00)*
R	−1.901 (0.03)*	−2.462 (0.01)*	−0.082 (0.47)	29.271 (0.40)	12.758 (0.99)	7.711 (0.00)*
Δln (*M*)	−19.954 (0.00)*	−15.588 (0.00)*	−20.591 (0.00)*	334.51 (0.00)*	359.55 (0.00)*	1.769 (0.04)*
Δln (*Y*)	−8.724 (0.00)*	−6.121 (0.00)*	−11.206 (0.00)*	176.380 (0.00)*	228.998 (0.00)*	1.112 (0.13)
Δ*R*	−15.630 (0.00)*	−12.781 (0.00)*	−13.682 (0.00)*	218.139 (0.00)*	242.821 (0.00)*	0.930 (0.18)

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) which is due to Maddala and Wu (1999) and Hadri (2000). In Hadri the null is that the variable is stationary.
Probability values are reported in the parentheses.
* and **Denotes the rejection of the null at 5% and 10% levels, respectively. For a discussion of these tests see Baltagi (2005) and Pesaran and Breitung (2005).

stable, central banks should use money supply as the monetary policy instrument. Using the rate of interest as the policy instrument will only accentuate instability.[4] 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.

The objectives of this paper are twofold. First, we examine, with the Pedroni (2000, 2004) Fully Modified Ordinary Least Squares (FMOLS) panel data methods, if there is a meaningful long run relationship between the demand for money and its determinants for a panel of 14 Asian countries. For comparisons we have also used two other alternative panel estimation methods of Mark and Sul (2003) and Breitung (2006).[5] Second, we will examine if there has been a structural change in this relationship leading to instability after financial reforms because this has implications for the choice of monetary policy instruments.

The second objective is difficult to test. However, we shall proceed as follows. In comparison to testing for unit roots in a variable with structural breaks, there are only a few works on structural breaks in the panel data cointegrating equations. Banerjee and Carrion-i-Silvestre (2006), BC hereafter, is one such recent and influential work. BC's method has some limitations, from an applied perspective, because they assume a single structural break at the beginning or in the middle or towards the end of the sample period. Consequently, it is not possible to determine the break date endogenously and estimate the parameters of cointegrating equations for the pre and post break samples. BC's main objective seems to be to show that their technique has more power than Pedroni's (2004) in which there are no structural breaks. Therefore, BC's method is especially useful if the Pedroni methods fail to yield plausible cointegrating equations. In another recent study Westerlund (2006) has developed a method to test for breaks in the deterministic components i.e., intercepts and trends. However, this has a limited use for our purpose because we are interested in the changes of the slope parameters. Furthermore, this method needs a large time series dimension and especially useful for quarterly and monthly data; see also Bagnani (2009) who notes similar limitations.

If financial reforms have been effective, it is to be expected that there would be a structural break in the cointegrating equation after the mid 1980s because these reforms have been implemented in the Asian countries after such reforms were implemented in the developed countries. From the demand for money perspective there should be some improved economies of scale meaning that income elasticity should show a decline and an improvement in the responsiveness of money demand to changes in the rate of interest because of more market based interest rate policies and improved capital mobility. There are no tests for the temporal instability of the panel data cointegrating equations which are similar to the popular CUSUM and CUSUMSQ tests in the country specific time series models. Furthermore, strictly speaking, the CUSUM tests are not tests for the temporal stability of the cointegrating equation because the long run money demand is a derived relationship and unobservable.[6] Therefore, one may hypothesize that if the long run demand for money has become unstable due to financial reforms, estimates of the cointegrating parameters, after the structural break, will be less robust or may yield implausible estimates or there is no cointegration between the variables. Consequently, we can only make plausible conjectures about the effects of financial reforms on the structure of the demand for money and its stability in the panel data methods. For this purpose it is necessary to estimate the demand for money for the sub-samples with observations before and after the reforms. However, it is difficult to select a date for the structural break because financial reforms were not introduced by all the Asian countries at the same time and with the same intensity.

While many East Asian countries have liberalized their financial markets from the early 1980s, the South Asian countries were late starters and delayed reforms until the early 1990s. Furthermore, reforms seem to have been introduced without considering the adequacy of the existing banking laws. Consequently the East Asian countries had a major financial crisis during 1997–1998. On the other hand in countries like India several non bank financial intermediaries, known as chit-funds, were established. They have mobilized large amounts of deposits but many have become insolvent and bankrupt due to the inadequacies in the Indian banking laws. Therefore, a single break date might be somewhat restrictive.

With this perspective, the outline of this paper is as follows. Section 2 briefly discusses the data and presents results for unit root and cointegration tests and estimates of the cointegrating equations for the entire sample period of 1970–2005 with the Pedroni (2004) FMOLS. For comparisons we also report estimates with the dynamic

[4] Poole's results are based on the instability in the IS and LM relations. However, instability in the demand for money is the major cause of instability in LM.

[5] These are known as the first generation panel data methods and assume homogeneity across the cross section units. The second generation panel methods allow for heterogeneity and are beyond the scope of this paper.

[6] What is tested with the stability tests like the CUSUMS is the stability of the parameters of the short run dynamic coefficients in the ARDL terms and the adjustment coefficient of the lagged error correction term (ECM). To test for the stability of the long run demand for money it is necessary first to estimate the cointegrating equation, for example, with the Gregory and Hansen (1992) method which allows for breaks. The lagged ECM from this can be used to estimate the short run dynamic adjustment equation. In the second stage, CUSUM and CUSUMSQ tests may be applied to test its stability. However, it is necessary for further developments for determining in panel data methods and BC note these limitations. Therefore, our aforesaid procedure should be interpreted with caution. Mark and Sul (2003) summarize the observed changes in the parameter estimates and in particular the decline in the income elasticity of the US demand for money. This indicates that the structure of the long run demand for money is somewhat susceptible to structural changes.

Table 2
Panel cointegration tests 1970–2005.

Test statistic	Fixed effects	Random effects
Panel ν- statistic	1.466	−0.269
Panel σ- statistic	−2.648*	−1.801**
Panel $\rho\rho$- statistic	−3.633*	−4.122*
Panel ADF-statistic	−2.176*	−2.888*
Group σ- statistic	−3.128*	−1.278
Group $\rho\rho$- statistic	−5.048*	−4.201*
Group ADF- statistic	−4.239*	−3.191*

Notes: The test statistics are distributed as $N(0,1)$.
* and **Denotes significance, respectively, at 5% and 10% levels.

Table 3
Estimates of the cointegration coefficients 1970–2005 dependent variable: ln(M).

	ln(Y)	R	ln(Y)	R
	Fixed effects		Random effects	
Pedroni	1.14* (20.84)	−0.02* (−5.60)	0.94* (79.98)	−0.01* (−7.74)
Mark and Sul	0.99* (32.00)	−0.01* (−2.75)	0.97* (19.88)	−0.01* (−2.75)
Breitung	0.96* (60.19)	−0.01* (−5.24)		

Notes: t-Ratios are in the parentheses and * indicates significance at the 5% level.

ordinary least squares method (DOLS) of Mark and Sul (2003) and a simple two-step procedure of Breitung (2006).[7] Both methods claim that they have better finite sample properties than Pedroni's. However, Breitung also claims that his technique gives less biased estimates in finite samples. Section 3 reports results of the estimated cointegrating parameters for two sets of sub-samples to determine if financial reforms had the expected effects on the parameters of the post-reform sub-samples. The selected break dates are 1985 and 1995. Finally Section 4 summarizes our findings, their policy implications and limitations.

2. Unit root and cointegration tests

Our panel data consists of 14 Asian countries (N=1....14) for the period 1970 to 2005 (T=1....36) The selected countries are Bangladesh (BGD), Indonesia (IDN), India (IND), Iran (IRN), Korea (KOR), Malaysia (MYS), Myanmar (MYAN), Nepal (NPL), the Philippines (PHL), Pakistan (PAK), Papua New Guinea (PNG), Singapore (SGP), Sri Lanka (LKA) and Thailand (THA).[8] Definitions of the variables and sources of data are in the Appendix A.

Results of the panel unit root tests, which are commonly used for non-stationary panel variables, are in Table 1. These tests give somewhat mixed results for ln(M). The Breitung test in which is the null that the variable is non-stationary is not rejected at the 5% level. However, in the LLC, IPS, ADF and PP tests in which the null is the same accept the null at only the 1% level. In the Hadri test the null is that the variable is stationary and it is rejected at the 5% level. For ln(Y), with the exception of the Breitung test, all other tests show that it is a non-stationary variable. For R, the null is rejected by the LLC and Breitung tests. All other tests indicate that R is non-stationary. In contrast, that the first differences of these three variables are stationary are not rejected by all the tests. Therefore, it is reasonable to conclude that that these variables are by and large $I(1)$ in their levels.

The standard specification for the demand for money in many cointegration studies is [9]:

$$lnM_{it} = \alpha_i + \beta_{it} lnY_{it} + \gamma_{it} R_{it} + \varepsilon_{it} \tag{1}$$

where ln$M$ is the log of real money ($M1$), ln$Y$ is the log of real GDP and R is the nominal short term rate of interest.

Test results for cointegration between the 3 variables of Eq. (1) are in Table 2. The majority of the reported 7 tests show that there is cointegration between these variables at the 5% level. Only the panel ν and group σ test statistics in the random effects model and panel ν statistic in the fixed effects model are insignificant at the 5% level and the rest are significant rejecting the null of no cointegration. Of these 7 tests the two ADF tests have more power against the null and they reject conclusively the null of no cointegration. Therefore, it can be concluded that the variables in Eq. (1) are cointegrated and a long run money demand function exists for the group as a whole and the members of the panel.

Table 3 gives the estimated panel group cointegrating parameters, for the fixed and random effects models, with the Pedroni FMOLS, Mark and Sul's DOLS and Breitung's two-step methods.[10] Since the panel group estimates are important for our main discussion, estimates of individual country cointegrating parameters are relegated to the Appendix A.

Estimates of income elasticity and semi-interest elasticity differ only marginally in these three methods and all are significant at the 5% level. Coefficient of the rate of interest has the expected negative sign and income elasticity is very close to unity in all the estimates. From the t-ratios in the table it is hard to admit that the Mark–Sul and Breitung methods are conclusively more efficient than the Pedroni method. However, in comparison to the Pedroni and Mark and Sul methods that assume a single cointegration equation, Breitung method is based on the systems method and allows for the existence of multiple cointegrating equations. While these alternative methods may be theoretically more efficient in finite samples, each method may perform differently depending on the estimated relationship and data. On the basis of the above estimates we may conclude that

[7] Mark and Sul's (2003) DOLS and Breitung's (2006) two-step method differ in their treatment of the intercept, trend and variables that influence dynamic adjustments in estimating the cointegrating equations. Collectively these variables are called nuisance variables. However, the common objective of all the three methods is to estimate unbiased and efficient parameters, especially in finite samples. There is no difference in their asymptotic properties; see Breitung's excellent introduction to the literature and his criticisms of FMOLS. Although Mark and Sul's DOLS and Breitung's two-step procedure claim that their methods are more efficient in finite samples, we take the view that when the real world data are used there seems to be no clear cut result to show that one is unequivocally better than the other. Therefore, it seems better to use all the three methods in the applied works because, in finite samples, efficiency may also depend on the estimated relationships, their specifications and the quality of data. Pedroni's methods are simpler to implement with popular software packages like RATS, EViews 6 and STATA. Some knowledge of and experience with GAUSS is necessary to implement the other alternatives. Dreger and Roffia (2007) briefly discuss the relative merits of these three methods. In their estimates of the demand for money with a panel of 10 countries (8 Central and Eastern European and 2 Mediterranean countries), efficiency of the Pedroni method was found to be as good as or even a shade better than Mark-Sul and Breitung methods. For an excellent exposition of panel data methods see Baltagi (2005) which is widely cited, although the chapter with non-stationary variables needs an update. An excellent exposition for the beginners of panel data methods with non-stationary variables is Murthy (2007). Coakley, Fuertes and Spagnolo (2003, 2004) are also very useful to understanding several concepts and to estimate group mean coefficients with simple OLS.

[8] Originally we have included Hong Kong but due to diverse data sources we could not get plausible estimates. The income elasticity for Hong Kong was found to be −2.5 and is not unexpected in panel estimates. In Mark and Sul's (2003) estimates of the demand for money, income elasticity for Norway was 2.64 and for New Zealand −1.23. Hong Kong was ignored since the data are not reliable.

[9] Additional variables like the inflation rate and/or exchange rate are added in some empirical works; see Bahmani-Oskooee and Rehman (2005). We did not include these variables because unit root tests showed that inflation is a stationary variable and foreign exchange holding is not a practical option in many Asian countries.

[10] The main differences between the three estimation methods can broadly be explained by comparing them with the well known methods used to estimate cointegrating equations with the country specific time series data. Pedroni's method uses the Phillips–Hansen FMOLS and Mark and Sul use the Stock–Watson DOLS. Breitung's method is different. In the first stage he uses the Johansen maximum likelihood method. The second stage equation is estimated with OLS with the polled results from the first stage with the constraint that the parameters of the cointegrating equation are the same in all the countries. The GAUSS code of Breitung does not report the first stage country cointegration parameters. We have treated Breitung estimates as if they are fixed effects estimates in our results tables.

Table 4
Estimates of the sub-period cointegration coefficients dependent variable: log(M).

	ADF for cointegration	Log(Y)	R	log(Y)	R
		Fixed effects		Random effects	
Pedroni 1970–1985	−2.96* (P) −3.13*(G)	0.82* (12.84)	−0.02* (−8.21)	0.77* (30.68)	−0.01* (−5.92)
Pedroni 1986–2005	−2.25* (P) −3.19*(G)	1.38* (11.72)	−0.01 (−1.11)	1.03* (54.46)	−0.01* (−4.09)
Mark and Sul 1970–1985		0.96* (4.68)	−0.03 (−0.13)	0.94* (5.91)	−0.05 (−0.29)
Mark and Sul 1986–2005		0.83* (10.49)	0.10** (1.66)	0.78* (7.26)	0.12* (2.39)
Breitung 1970–1985		0.86* (19.89)	−0.02* (13.18)		
Breitung 1986–2005		1.00* (30.20)	−0.02* (−2.33)		
Pedroni 1970–1995	−4.36(P)* −5.39(G)*	0.87* (52.65)	−0.01* (−5.30)	0.97* (19.40)	−0.01* (−3.99)
Pedroni 1996–2005	−3.46(P)* −3.73(G)*	0.81* (17.92)	−0.01* (−5.00)	1.47* (11.08)	−0.02* (−3.14)
Breitung 1970–1995		0.94* (35.41)	−0.01* (2.76)		
Breitung 1996−2005		0.91* (10.36)	−0.01* (2.73)		

Notes: See notes for Table 3. (*P*) is panel ADF and (*G*) is group ADF test statistic for the fixed coefficient models. Estimates for the panel members are in the Appendix A.

income elasticity is about unity and money demand is responsive to changes in the rate of interest albeit this response is small.[11]

3. Effects of financial reforms

Financial reforms have been implemented globally from the mid 1970s although it is hard to say that all countries have implemented these reforms with the same vigor and at the same time. For example Singapore and Hong Kong have liberalized their economies much earlier than other East Asian countries like Malaysia, Thailand, Korea and Indonesia. In comparison, liberalization policies have started rather late from the early 1990s in the South Asian countries e.g., India, Pakistan and Sri Lanka etc. Therefore, it is hard to select a common break date for all the countries in our panel. The UN ESCAP Poverty Division (1997) briefly reviewed financial reforms in four major East Asian countries and their adequacy. In Malaysia these reforms started after some financial crises during the early 1980s. In Korea, Thailand and Indonesia reforms started earlier to support mainly their export industries. Therefore, 1985 is selected as a break date to capture some of these effects due to reforms. In India reforms started in the early 1990s due to its precarious foreign exchange reserves. Other South Asian countries followed India. To capture these effects a break date of 1995 seems to be reasonable.

Before further discussion, it would be useful to take an overview of what is expected from these sub-sample estimates. Firstly, we are looking for some evidence on whether financial reforms had any significant effects. If they have been effective, it is to be expected that there will be evidence for some economies of scale in the use of *M1* and also the response of the demand for money to the rate of interest will improve because of more market based interest rate policies. Therefore, it is to be expected in the second set of sub-samples income elasticity will show a decline and the absolute value of the interest rate coefficient will increase and/or become significant if it was insignificant in the pre-reforms sample. Second, if reforms have created substantial number of near monies and if this is a continuous process, this may lead to instability in the demand for money. This should be reflected in the second set of sub-samples as lack of a well defined long run relationship between money and its determinants i.e., cointegration tests might show that there is no cointegration. Furthermore, even if these tests reject the null of no cointegration, the estimated parameters may have large standard errors to make them insignificant.

In the two sets of sub-samples, the null of no cointegration is rejected by the majority of the cointegration tests and the details are reported in the Appendix in Table A3. The more powerful ADF test statistics are reported in the rows for the Pedroni tests in Table 4. These are significant at the 5% level, whether the model is a fixed (reported) or random (not reported) coefficients model. The only exception is the panel ADF test statistic for the random coefficients model for the sub-period 1996–2005; see Table A.3. By and large there is no strong evidence that there is no cointegration in the two sets of sub-sample periods.

Estimates of the cointegrating parameters for two sets of sub-samples, with break dates in 1985 and 1995 are in Table 4. The Pedroni and Breitung estimates, with a break date in 1985, imply that there is no evidence to support that reforms had any significant expected scale and interest rate effects. The estimates for the first set of sub-samples show that in the post-reforms period scale economies might have actually decreased and the response to the rate of interest has declined or remained the same. However, estimates with the Mark and Sul method indicate that scale economies might have shown some improvement in the post-reform sample. Income elasticity has declined from 0.96 from the pre-reforms period to 0.83 in the post reforms period. But the sign of the coefficient of the rate of interest has changed from an insignificant −0.03 in the pre-reform period to 0.11 which is significant at the 10% level. Therefore, it is somewhat difficult to choose between the evidence from these alternative methods. Since both in the Pedroni and Breitung methods the coefficients have the expected signs and in the Mark and Sul method interest rate became positive and also failed in our subsequent estimation with 1995 as the break date, we prefer the estimates with the Pedroni and Breitung methods.[12] On this basis, financial reforms did not seem to have the expected scale and rate of interest effects. This might be, as some commentators have observed, due to the limited nature of these reforms and/or 1985 is too early to expect that reforms have worked. That the reforms were not strong enough and inadequate was corroborated by the East Asian financial crisis in 1997–1998, by a panel of advisors to the Ministry of Finance of Japan and Mundle (2004).[13]

In the country specific estimates (not reported to conserve space but can be obtained from us) there is some evidence to conclude that financial reforms might have been effective in India where the income

[11] Pedroni's methods gave the highest and lowest point estimates of income elasticity which are 1.14 and 0.94. Their 1.96 times standard errors limits range from 1.25 and 1.03 for the first and 0.96 to 0.92 for the lowest value. Strictly speaking income elasticity could be slightly less than unity by about 4% which is negligible.

[12] Mark and Sul give the options of including an intercept and country specific time trend. Their second equation with the intercept is the fixed effects model and the third equation with both an intercept and trend is their random effects model. In our subsequent estimates, with 1995 as the break date, neither of these options produced any sensible results for the sub-samples. Therefore, these are estimated with their first equation without an intercept and trend. This seems to be close to the fixed coefficients model in the panel data methods.

[13] In 2003 a panel of advisors to the Ministry of Finance of Japan observed that "Asian countries need to bolster reform of their financial sectors to achieve economic growth and prevent a financial crisis like the one in 1997–1998"; an abstract of this report is at http://findarticles.com/p/articles/mi_m0WDP/is_2002_July_8/ai_88685527. Mundle (2004) commenting on the inadequacy of the financial reforms that has caused the 1997–1998 crisis says that "The Asian financial crisis of 1997 was characterized by a combination of several inter-related processes: appreciating dollar, pegged exchange rates, declining exports, arbitrage of domestic and international interest rate differentials in liberalized capital account regimes, reckless financial intermediation, asset price collapse, debt defaults, etc." A comprehensive survey see Woo, Parker and Sachs (1997).

elasticity has declined from 1.29 to 1.02 and the coefficient of the rate of interest rate, which was insignificant during 1970–1985 has become significant with a value of −0.04. However, it is well known that India started its reforms after 1991. At best, we may conclude from this evidence that the reforms implemented by these countries have been not significantly effective by 1985.

Estimates for the sub-samples with 1995 as the break date have also given very similar inconclusive results although the scale economies have increased by a negligible amount of about 3% in the Pedroni and Breitung estimates. The response to the rate of interest remained the same and significant. On the other hand the Mark and Sul method did not yield estimates for these two sub-samples because the inverse matrices for estimating the coefficients were not positively definite. This is possible if there is more than one cointegrating equation. Therefore, the Breitung systems method seems to be preferable for these sub-samples.[14] On the basis of these results we conclude that there is no structural break and instability in the demand for money in our panel of 14 Asian countries. If the long run demand for money for the individual countries shows instability, for which there is no evidence as yet, financial reforms are not a major cause for such instability. This conclusion is consistent with Bahmani-Oskooee and Rehman's (2005) findings that money demand functions have been fairly stable in many Asian countries and with the recent findings of Rao and Singh (2005) for India and Sumner (2009) for Thailand. This implies that using the interest rate as the monetary policy instrument by the central banks of Asia is somewhat premature, inappropriate and may actually accentuate economic instability.

4. Conclusions and limitations

This paper has used 3 alternative panel data methods of Pedroni, Mark and Sul and Breitung, to estimate the long run demand for money for a panel of 14 Asian countries. Our results show that these 3 methods yield similar parameter estimates and with similar efficiency. However, this conclusion cannot be generalized because their efficiency may also depend on other empirical considerations. Therefore, it is desirable to use these 3 alternative methods in applied works.

Estimates for the entire sample period of 1970 to 2005 showed that income elasticity of demand for money is about unity and demand for money responds negatively to variations in the short term rate of interest, albeit by a small amount. This framework is extended to test if the financial reforms in these countries have had any significant effects. Our sub-sample estimates show that reforms do not seem to have had any significant effects so far. This may be due to various factors like the difficulties to effectively implement reforms and/or due to the inadequate nature of such reforms.

An implication of our results is that financial reforms are not a major contributor to the instability (which may be non-existent) in the demand for money. Further, there is no evidence to say that the long run demand for money has become unstable because cointegration tests for the sub-samples reject the null of no cointegration. Therefore, central banks of these countries should use money supply as their monetary policy instrument to achieve their policy targets like moderate inflation rates and overall economic stability. Imitating the central banks in the advanced countries may actually lead to more instability in the economy.

Needless to say there are many limitations in this paper. Firstly, while the results for the entire sample period of 1970–2005 are impressive and close in the three alternative estimation methods, estimates for the individual countries are not always impressive. For some countries like Sri Lanka income elasticity is as high as 3% and for Nepal it is as low as 0.12% and insignificant; see Table A.1. Such results are not unusual in panel data estimates based on the first generation unit roots and cointegration tests. Therefore, there is a need to use the second generation panel data methods; see Pesaran and Breitung (2005) for a survey of the literature. Second, our choice of the break dates is somewhat arbitrary, but then as yet there is no satisfactory panel data method to estimate the break dates endogenously. Nevertheless we hope that our paper will stimulate further theoretical and empirical work to make panel data methods popular in applied work. For this purpose it seems necessary for further theoretical developments with endogenous structural breaks in the panel data methods. Finally, our results showed that the Pedroni methods are not any less efficient than Mark and Sul and Breitung methods. Therefore, further theoretical and empirical work, preferably with data from the real world, seems necessary to evaluate their finite sample properties.

[14] For other advantages of the Breitung method see Pesaran and Breitung (2005). Breitung claims that the bias of estimates in finite samples with his systems method is much less than in the FMOLS and DOLS methods.

Appendix A

Pedroni's country specific cointegration coefficients 1970–2005.

	ln (Y)	R	Ln (Y)	R
	Fixed effects model		Random effects model	
Bangladesh	2.20 (3.98)*	0.00 (0.01)	0.81 (7.63)*	−0.02 (1.09)
Indonesia	1.11 (11.10)*	−0.00 (1.27)	1.02 (56.23)*	−0.00 (2.17)*
India	1.54 (3.74)*	0.03 (1.53)	1.05 (24.60)*	−0.04 (3.63)*
Iran	1.05 (3.83)*	−0.08 (1.62)	0.24 (0.66)	0.09 (0.96)
Korea	0.38 (4.03)*	0.01 (2.63)*	0.82 (23.05)*	−0.00 (0.11)
Malaysia	1.45 (14.92)*	−0.01 (1.32)	1.12 (43.36)*	−0.02 (2.27)*
Myanmar	0.66 (4.93)*	−0.02 (3.48)*	0.87 (12.25)*	0.00 (0.25)
Nepal	0.12 (0.37)	−0.03 (4.00)*	1.55 (14.92)*	0.01 (1.12)
Pakistan	1.79 (5.89)*	−0.01 (1.32)	1.00 (20.38)*	−0.00 (0.19)
Philippines	0.65 (3.70)*	−0.05 (3.72)*	1.22 (13.14)*	−0.04 (6.02)*
Papua New Guinea	0.85 (5.14)*	−0.03 (1.56)	1.07 (11.92)*	−0.06 (5.71)*
Singapore	0.74 (12.21)*	0.00 (0.20)	0.86 (30.47)*	−0.02 (1.84)**
Sri Lanka	3.12 (2.56)*	−0.05 (2.51)*	0.64 (13.24)*	0.01 (1.61)
Thailand	0.29 (1.59)	−0.06 (4.55)*	0.83 (27.42)*	−0.05 (9.87)*

Notes: The absolute t-ratios are reported in parentheses.
* and **Denotes significance at the 5% and 10% levels, respectively.

Panel cointegration tests for the pre-reforms sub-samples.

Test statistic	Fixed effects	Random effects
Panel υ- statistic		
1970–1985	0.240	−1.614
1970–1995	1.191	−0.691
Panel σ- statistic	−1.090	−0.364
1970–1985	−2.615*	−1.200
1970–1995		
Panel $\rho\rho$- statistic		
1970–1985	−4.762*	−6.192*
1970–1995	−4.700*	−5.177*
Panel ADF-statistic	−2.964*	−5.183*
1970–1985	−4.358*	−4.05400
1970–1995		
Group σ- statistic	−0.505	−1.917**
1970–1985	−1.980*	−0.373
1970–1995		
Group $\rho\rho$- statistic	−5.753*	−6.302*
1970–1985	−5.403*	−5.657*
1970–1995		
Group ADF- statistic	−3.135*	−5.761*
1970–1985	−5.388*	−5.190*
1970–1995		

Notes: The test statistics are distributed as $N(0,1)$. The critical values at 5% and 10% levels are 1.96 and 1.64, respectively.
* and **Denotes significance, respectively, at 5% and 10% levels.

Panel cointegration tests for the post-reforms sub-samples.

Test statistic	Fixed effects	Random effects
Panel υ- statistic		
1986–2005	1.069	−0.831
1996–2005	1.425	−0.841

Appendix A (*continued*)

Test statistic	Fixed effects	Random effects
Panel σ- statistic		
1986–2005	−1.297	−0.649
1996–2005	−0.079	−2.467*
Panel $\rho\rho$- statistic		
1986–2005	−3.000*	−4.700*
1996–2005	−3.431*	−0.754
Panel ADF-statistic		
1986–2005	−2.249*	−4.414*
1996–2005	−3.456*	−0.900
Group σ- statistic		
1986–2005	−0.268	0.956
1996–2005	−1.949**	3.806*
Group $\rho\rho$- statistic		
1986–2005	−3.212*	−4.348*
1996–2005	−3.234*	−0.444
Group ADF- statistic		
1986–2005	−3.187*	−5.180*
1996–2005	−3.726*	−3.297*

Notes: The test statistics are distributed as N(0,1). The critical value at 5% level is 1.96.
*Denotes significance at 5% level.
Cointegration test results with 1995 as the break date can be obtained from the authors.

Pedroni's country specific cointegration coefficients 1970–1985.

	ln (Y)	R	ln (Y)	R
	Fixed effects model		Random Effects Model	
Bangladesh	0.79 (0.57)	−0.06 (1.14)	−0.48 (0.98)	0.09 (2.33)*
Indonesia	1.24 (4.32)*	−0.01 (2.43)*	1.03 (30.42)*	−0.01 (5.73)*
India	1.45 (2.50)*	0.07 (2.30)*	1.29 (1.45)	−0.10 (0.76)
Iran	−0.26 (1.25)	−0.07 (1.67)**	0.19 (0.18)	0.05 (0.33)
Korea	0.10 (0.31)	0.01 (0.68)	0.70 (9.34)*	0.00 (0.10)
Malaysia	0.69 (3.45)*	0.03 (1.56)	0.85 (10.89)*	0.01 (0.94)
Myanmar	1.76 (7.59)*	−0.04 (4.71)*	1.19 (9.27)*	−0.05 (0.52)
Nepal	−0.25 (1.45)	−0.02 (3.44)*	1.55 (8.10)*	−0.05 (2.18)*
Pakistan	2.10 (6.95)*	−0.03 (3.72)*	0.86 (2.31)*	0.00 (0.02)
Philippines	1.28 (9.34)*	−0.05 (12.14)*	0.81 (8.36)*	−0.03 (7.25)*
Papua New Guinea	1.61 (13.11)*	−0.06 (3.68)*	0.58 (1.71)**	−0.06 (3.16)*
Singapore	0.86 (5.29)*	0.00 (0.28)	0.86 (13.26)*	−0.01 (0.44)
Sri Lanka	0.94 (1.47)	−0.02 (1.76)**	0.72 (6.49)*	−0.01 (1.66)**
Thailand	−0.84 (4.15)*	−0.01 (0.88)	0.57 (14.01)*	−0.02 (4.15)*

Notes: The absolute t-ratios are reported in parentheses.
* and **Denotes significance at the 5% and 10% levels, respectively.

Pedroni's country specific cointegration coefficients 1986–2005.

	ln (Y)	R	ln (Y/P)	R
	Fixed effects model		Random effects model	
Bangladesh	0.15 (0.21)	0.01 (1.15)	1.19 (7.43)*	−0.00 (0.25)
Indonesia	0.81 (6.49)*	−0.00 (0.78)	1.01 (29.99)*	−0.00 (0.68)
India	2.15 (6.27)*	−0.02 (1.14)	1.08 (21.09)*	−0.04 (5.16)*
Iran	4.16 (5.59)*	0.01 (0.29)	0.29 (2.56)*	0.24 (5.25)*
Korea	0.87 (2.37)*	0.01 (2.95)*	1.01 (10.46)*	0.00 (1.24)
Malaysia	1.87 (7.08)*	−0.02 (1.26)	1.23 (22.23)*	−0.01 (0.73)
Myanmar	0.33 (2.26)*	0.01 (0.99)	0.73 (11.59)*	0.02 (2.72)*
Nepal	0.56 (0.95)	−0.02 (1.34)	1.54 (14.71)*	0.02 (2.62)*
Pakistan	1.98 (4.76)*	−0.00 (0.31)	1.06 (5.22)*	0.00 (0.18)
Philippines	−0.64 (2.66)*	0.00 (0.17)	1.84 (14.89)*	−0.00 (0.84)
Papua New Guinea	0.35 (1.49)	−0.02 (0.79)	1.24 (5.61)*	−0.05 (2.88)*
Singapore	0.49 (5.17)*	0.01 (2.04)*	0.84 (36.19)*	−0.06 (8.19)*
Sri Lanka	6.07 (3.12)*	−0.01 (0.25)	0.37 (9.84)*	0.00 (0.11)
Thailand	0.21 (0.75)	−0.07 (5.86)*	0.92 (11.96)*	−0.05 (8.71)*

Notes: The absolute t-ratios are reported in parentheses.
*Denotes significance at the 5% level.
Country specific estimates for the sub-periods with 1995 as the break date can be obtained from the authors.

Data Appendix

Y = Real GDP at factor cost. Data are from (IFS-2005) and ADB database (2005).

R = The average of 1–3 years savings deposit rate. Data are from (IFS-2005) and ADB database (2005).

M = Real narrow money supply. Data are from (IFS-2005) and ADB database (2005).

Note:

All variables, except the rate of interest, are deflated with the GDP deflator and converted into natural logs.

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