A Panel Data Approach to the Demand for Money and the Effects of Financial Reforms in the Asian Countries

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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

Three panel data estimation methods are used to estimate the cointegrating equations for the demand for money (M1) in 14 developing Asian countries. Tests for the effects of financial reforms are made with estimates for two sub-samples of 1970-1985 and 1986-2005. Our results show that money demand functions in these Asian countries are stable and financial reforms have yet to have any significant effects. This implies that the central banks of these countries should use money supply, instead of the rate of interest, as the monetary policy instrument.

Keywords: Pedroni, Mark and Sul and Breitung methods, Demand for money, Asian countries, Effects of financial reforms and Choice of monetary policy instruments.

JEL: E41 and E58

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1. Introduction

Demand for money and its stability have received vast attention. 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 a stylized fact that demand for narrow and broad money have become temporally unstable in the developed countries after the financial reforms. This is attributed to the effects of deregulation of the financial markets which has increased competition in the financial markets, created additional money substitutes, increased the use of credits cards and electronic money transfers, increased liquidity of the fixed deposits and induced higher international capital mobility. Consequently many central banks in the developed countries have abandoned using money supply as a policy instrument because it is difficult to predict demand for money with temporally unstable functions. Furthermore, the Taylor rule has made more attractive the use of bank rate as the policy instrument by arguing that it will increase the built-in stability of the economy. Therefore, many central banks in the developed countries have abandoned using money supply and switched to the rate of interest as their monetary policy instrument since the mid 1980s. This switch is also consistent with Poole (1970) who showed that rate of interest should be targeted if demand for money is unstable.

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) showed that demand for money functions in several developing Asian countries, by and large, are stable.[2] According to Poole (1970) if demand for money is stable, central banks should target money supply. Targeting the rate of interest will only accentuate instability. Therefore, it is important to know if there are stable money

[2] The countries selected in this study are India, Indonesia, Malaysia, Pakistan, the Philippines, Singapore and Thailand. They find 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 for M1 are not well determined.

demand function in the developing countries because stability implies that targeting the rate of interest is inappropriate.

The objectives of this paper are twofold. First, we examine, with the Pedroni (2002) panel data methods, if there is a meaningful long run relationship between the demand money and its determinants for a group of selected Asian countries. Second, we examine if this relationship exists and stable after financial reforms because this has implications for the choice of monetary policy instruments.

The second objective is difficult to test. But because of its importance we proceed as follows. In comparison to testing for unit roots with structural breaks, there are only a few works on cointegration with structural breaks in panel data of which Banerjee and Carrion-i-Silvestre (2006), BC hereafter, is the most recent. 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 parameters of the cointegrating equations before and after the break date. BC's main objective seems to be to show that their technique has more power than Pedroni's (2004) without structural breaks. Therefore, BC's method is especially useful if the Pedroni methods fail to yield plausible cointegrating equations.

If financial reforms are effective, it is to be expected that there would be a structural break in the cointegrating equation after the mid 1980s with a decline in income elasticity due to improved economies of scale and an increase in the interest rate elasticity due to more market oriented interest rate policies and more capital mobility. It is difficult to test for the temporal instability of the cointegrating equation with tests similar to the CUSUM and CUSUMSQ tests used in the country specific time series models. Furthermore, strictly speaking, these are not tests for the temporal stability of the cointegrating equation because the long run money demand is a derived relationship and unobservable. Therefore, one may hypothesize that if the long run demand for money is unstable estimates of the cointegrating parameters, before and after the structural break, will be less robust and may yield implausible estimates or that there is no cointegration between the variables.[3] Consequently, we can only make plausible conjectures about these structural changes and instability in the long run demand for money with panel data. For this purpose it is necessary to estimate the demand for money for the sub-samples with observations before and after the reforms.

[3] What is tested with the stability tests, e.g., CUSUM and CUSUMSQ, is the stability of the parameters of the short run dynamic adjustment coefficients in the ARDL terms and the adjustment coefficient of the lagged error correction term. To test for the stability of the long run demand for money it seems necessary first to estimate the cointegrating equation, for example, with the Gregory and Hansen (1992) method to allow for a single break. The lagged error correction term from this estimate 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 in the estimation methods to estimate cointegrating equations with structural breaks with the dynamic adjustment terms and test its stability. BC note these limitations in the present developments. Therefore, the aforesaid procedure should be interpreted with caution. Mark and Sul (2003) elegantly summarize the observed changes in the parameter estimates and in particular the decline in the income elasticity of the US demand for money which indicate that the structure of the long run demand for money is more likely to change.

With this perspective, the outline of this paper is as follows. Section 2 briefly discusses the data and presents results of the unit root and cointegration tests and estimates of the cointegrating equations with the Pedroni (2004) method based on the fully modified ordinary least square (FMOLS). For comparisons we shall also report estimates with the dynamic ordinary least squares method (DOLS) of Mark and Sul (2003) and a simple two step procedure of Breitung (2006).[4] Both methods claim that they have better finite sample properties than Pedroni's. Section 3 contains results of the estimated cointegrating parameters for the sub-samples to determine if financial reforms had the expected effects on the parameters in the sub-samples. Finally Section 4 summarizes our findings, policy implications and limitations of this paper.

2. Estimates with Alternative Methods

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

[4] Alternatives to Pedroni's FMOLS are Mark and Sul's (2003) DOLS and Breitung's (2006) two-step method. They differ in their treatment of the intercept, trend and variables that influence dynamic adjustment in estimating the cointegrating equations of the panel members. Collectively these variables may be called nuisance variables. But the common objective of these alternative methods is to estimate unbiased and efficient parameters of the cointegrating equation and satisfy this objective asymptotically. Therefore, choice between them is based on their claimed finite sample properties and there is no clear cut result to show that one is better than the other. We take the view that it is better to use all the three methods because efficiency may also depend on the estimated relationship, specifications and 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 two alternatives. Dreger and Roffia (2007) briefly discuss, from an applied perspective, 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) efficiency of the Pedroni method was as good or even better than the Mark-Sul and Breitung methods. For an excellent exposition of panel data methods see Baltagi (2006) which is a classic now. An excellent exposition for the beginners in panel data methods is by Murthy (2007).

(LKA) and Thailand (THA).[5] Definitions of the variables and sources of data are in the appendix.

Results of the panel unit root tests, which are generally used in the empirical work with the non-stationary panel variables, are in Table 1. These tests give somewhat mixed results. While the Hadri and Breitung tests confirm at the 5% level that ln (M) is non-stationary, LLC, IPS and PP tests confirm this only at the 1% level. Except the Breitung test other tests confirm that ln (Y) is non-stationary and except LLC and Breitung tests other tests confirm that R is non-stationary. However, that the first differences of these variables are stationary is confirmed by all the tests at the 5% level and it is reasonable to conclude that that these variables are by and large I(1) in their levels.

Table 1. Panel Unit Root Tests 1970-2005

Series	LLC	Breitung- *t*	IPS- *W*	ADF	PP	Hadri
ln (*M*)	-1.977 (0.02)*	2.461 (0.99)	-2.061 (0.02)*	52.132 (0.003)*	54.082 (0.002)*	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 (0.007)*	-0.082 (0.47)	29.271 (0.40)	12.758 (0.99)	7.711 (0.00)*
? ln (*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)*
? ln (*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)
? ln *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)

[5] Originally we included Hon 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 data. For example in Mark and Sul's (2003) estimates of the demand for money, income elasticity for Norway was high at 2.64 and as implausible as -1.23 for New Zealand. We removed Hong Kong from our sample mainly because the data are not reliable.

Notes: The tests are: Levin, Lin and Chu (2002) (LLC), Breitung (2000), Im, Pesaran and Shin (2003) (IPS), ADF Fisher Chi-Square (ADF), PP Fisher Chi-Square (PP), and Hadri (2000). In the 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.

The standard specification for the demand for money in many cointegration studies is[6]:

$$\ln M_{it} = \alpha_i + \beta_{it} \ln Y_{it} + \gamma_{it} R_{it} + \varepsilon_{it} \qquad (1)$$

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

Test results for cointegration between the 3 variables in (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 ? and group s test statistics in the random effects model and panel ? 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 (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, with the fixed and random effects, with the Pedroni FMOLS, Mark and Sul's DOLS and Breitung's two-step methods. Estimates of individual country cointegrating parameters are in the appendix since the panel group estimates are important for our discussion.

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. While these alternative methods may

[6] Additional variables like the inflation rate and/or exchange rate 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.

Table 2. The Pedroni Panel Cointegration Tests 1970-2005

Test Statistic	Fixed Effects	Random Effects
Panel ?- statistic	1.466	-0.269
Panel *s* - statistic	-2.648*	-1.801**
Panel ??- statistic	-3.633*	-4.122*
Panel ADF-statistic	-2.176*	-2.888*
Group *s* - statistic	-3.128*	-1.278
Group ??- 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 Sue	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.

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 income elasticity is about unity and money demand is responsive to changes in the rate of interest albeit this response is small.[7]

3. Effects of Financial Reforms

Financial reforms have been implemented globally from the early 1980s although it is hard to say that all countries have implemented these reforms with the same vigor and at the same time. Therefore, we have arbitrarily selected 1985 as the break point and re-estimated money demand functions for the periods 1970 to 1985 and 1986 to 2005 and these are in Table 4.

Before any 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 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 sub-sample that income elasticity will show a decline and semi-interest rate elasticity may increase and/or become significant if it was insignificant in the pre-reforms sample. Second, if reforms have created 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 sub-sample 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 become implausible and/or their standard errors will be large to make them insignificant.

In the sub-samples, the null of no cointegration is rejected by the majority of the cointegration tests and these are reported in the appendix. The more powerful ADF test statistics are reported in the rows for the Pedroni tests in Table 4. Since these are significant at the 5 % level, the null that there is no long run demand for money in each sub-sample

[7] 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.

should be rejected. Estimates of the cointegrating parameters for the sub-samples, with the 3 methods, are also shown in Table 4.

Estimates with the Pedroni and Breitung methods imply that there is hardly any evidence to show that financial reforms had any significant effects. Income elasticity in both sub-samples is close and in fact it seems to have increased somewhat in the Pedroni and Breitung estimates. Coefficients of the rate of interest have remained negative, significant and virtually unchanged. The increase in the income elasticity may be due to the increasing monetization in many developing countries in our sample. However, estimates with the Mark-Sul method indicate that after the reforms there is a small improvement in the scale economies but the coefficient of the rate of interest is positive in the reforms period and insignificant in the pre-reforms period. It is hard to derive any firm conclusions on the effects of financial reforms on the basis of these estimates. 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 had the expected effects only in India where income 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. At best, we may conclude that the reforms implemented by these countries have been not yet fully effective or that they are not strong enough.

Therefore, if the long run demand for money in the individual countries shows instability, financial reforms are not the major cause and using the interest rate as the monetary policy instrument is an inappropriate. Central banks in these countries should use money supply as their monetary policy instrument because there is no convincing evidence that the long run relationship between money and its determinants has significantly changed and/or unstable due to the financial reforms. Our findings are also consistent with Bahmani-Oskooee and Rehman's (2005) that demand for money has been fairly stable in many Asian countries.

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)

Notes: see notes for Table 3. (P) is panel ADF and (G) is group ADF test statistic. Estimates for the panel members are in the appendix.

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 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 undertaken by 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 mild nature of such reforms.

An implication of our results is that financial reforms are not a major contributor to the instability in 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. Imitating the central banks' policies in the advanced countries may actually lead to more instability in the economy.

Needless to say there are several limitations in this paper. Firstly, although the results for the entire panels are impressive, 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. Although these are not unusual in panel data estimates, further attention to the quality of data seems to be necessary or it may be necessary to include some missing variables into the specification. However, it is difficult to introduce country specific special factors in the panel data methods. Second, our choice of the break date is somewhat arbitrary, but then as yet there is no satisfactory panel data method to estimate the break date endogenously. Nevertheless we hope that our paper will stimulate further theoretical and empirical work to make panel data methods popular in many applied works.

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:

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

Appendix for Tables

Table A.1. 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.

Table A.2. Pedroni's Panel Cointegration Tests 1970-1985

Test Statistic	Fixed Effects	Random Effects
Panel ?- statistic	0.240	-1.614
Panel s- statistic	-1.090	0.364
Panel ??- statistic	-4.762*	-6.192*
Panel ADF-statistic	-2.964*	-5.183*
Group s- statistic	0.505	1.917**
Group ??- statistic	-5.753*	-6.302*
Group ADF- statistic	-3.135*	-5.761*

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.

Table A.3. The Pedroni Panel Cointegration Tests 1986-2005

Test Statistic	Fixed Effects	Random Effects
Panel ?- statistic	1.069	0.831
Panel s- statistic	-1.297	-0.649
Panel ??- statistic	-3.000*	-4.700*
Panel ADF-statistic	-2.249*	-4.414*
Group s- statistic	-0.268	0.956
Group ??- statistic	-3.212*	-4.348*
Group ADF- statistic	-3.187*	-5.180*

Notes: The test statistics are distributed as $N(0,1)$. The critical values at 5% level is 1.96. * denotes significance at 5% level.

Table A.4. 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.

Table A.5. 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.

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