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Money demand function with asymmetric adjustment: Evidence on Brazil, Russia, India and China (BRICs)

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The long-run equilibrium relationship for the money demand function in the BRICs is investigated by the asymmetrical TAR and M-TAR cointegration tests developed by Enders and Granger (1998), Enders and Siklos (2001). Empirical results indicate that real M2 money balance, real GDP, real exchange rate and deposit rate have a long term relationship under some specific threshold value. Furthermore, we apply asymmetrical error-correction models to test if the money demand of the BRICs exist any nonlinear forms which will be compared with symmetrical error-correction models. Therefore, we find that M2 money demand in the BRICs support the hypothesis of an asymmetrical error correction process and provide better interpretation of macroeconomic meanings in the demand for money.

Key words: Money demand function, threshold autoregressive model, asymmetry adjustment, BRICs.

INTRODUCTION

A stable money demand function may be considered very important for conducting monetary policy, thus the money demand function has long been a cornerstone in macro-economic modeling. Academic researchers continue to search for a specification of the money demand function that gives a reliable long-run equilibrium relationship with other macroeconomic variables. Econometric estimates of money demand function abound in the developed and developing countries and most studies of the demand for money focus on developed countries. By comparison, emerging economies in general and transition economies in particular have received less attention in the literature. In this study, taking a fresh look at this function of the BRICs that includes China, India, Brazil, and Russia. Goldman Sachs (Wilson and Purushothaman, 2003) argue that the BRICs economies could become a much larger force in the world economy over the next 50 years, and predicts China and India, respectively, to be the dominant global suppliers of manufactured goods and services while Brazil and Russia would become similarly dominant as suppliers of raw materials. By 2025 they could account for over half the size of the G6. Thus, the BRICs have the potential to form a powerful economic bloc to the exclusion of the modern-day G6 status. These countries are forecast to encompass over 40% of the world's population and hold a combined GDP of 14.051 trillion. On almost every scale, they would be the largest entity on the global stage. Rao and Singh (2006) estimate the demand for narrow money in India and evaluate its robustness, thus they find that there is a stable demand for money for almost half a century. Deng and Liu (1999) use the cointegration and error-correction model to formulate the function of money demand and merge the short-run and long-run equations to give forecasts over different horizons in China. Austin, Ward, and Dalziel (2007) investigate nonlinearities in the demand for money in China that would suggest a threshold point for inflation materially entering into the decisions of Chinese house-holds and firms. Bahmani-Oskooee and Barry (2000) find that the demand for money in Russia which includes income, inflation rate and exchange rate variables were unstable in the 1990s. Harrison and Vymyatnina (2005) test the stability of long-run and short-run demand for money in Russia using M1 and M2 money aggregates and find some evidence of stability, but the adjustment

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lag is relatively long and money demand functions demonstrate signs of instability over the period.

The literature on money demand function estimation is long-standing and extensive. Most of these literatures are concerned with the existence of a stable money demand function. Notable references are Friedman and Schwartz (1991), McNown and Wallace (1992), Stock and Watson (1993), Ball (2001), and Anderson and Rasche (2001). Friedman and Schwartz (1991) have argued that the underlying characteristics of the money demand function rarely change over a long period of time. In the long run, the money demand function depends mainly on macroeconomic variables, such as interest rate, real income, and, in an open economy, exchange rates, may be cointegrated and have a stable long-run equilibrium relation. One way of examining the long-run equilibrium relationship of a money demand function is to test for a cointegration relationship. If a linear combination of non-stationary variables including real money balances, real income, and interest rates is stationary, the variables have a long-run equilibrium relationship. Using cointegration approaches, many studies have investigated the long-run equilibrium relationship of money demand functions. For example, Hafer and Jansen (1991), Baba et al. (1992), MacDonald and Taylor (1992), and Arize (1994), who investigated the stability of the US money demand function, used methods proposed by Engle and Granger (1987) and Johansen (1988), whose approaches assume that the adjustment process toward equilibrium is symmetric. In contrast, Friedman and Kuttner (1992) and Miyao (1996), employing these standard approaches, found the instability of the United State money demand function.

However, most of the research addressing the issue of equilibrium has not taken into account the asymmetric properties of the adjustment process in money demand. Asymmetry has been an important property in recent macroeconomic analysis, with a large number of studies providing evidence of the asymmetric adjustment of macroeconomic variables. Muscatelli and Spinelli (1996) and Ericsson et al. (1998) include a cubic error-correction term as a regressor in their study of the money demand function in Italy and in the United Kingdom, respectively. They find that the nonlinear error-correction specification better describes the short-run dynamics and improves the overall goodness-of-fit. For instance, the variations in money are more volatile in an economic downturn than in an upswing. Therefore, it is necessary to analyze the long-run equilibrium relationship in money demand by a cointegration test assuming asymmetric adjustment.

Threshold cointegration methods are being increasingly employed to analyze economic and financial data, and are especially useful in the study of time series which are characterized by asymmetric adjustment. The aim of this study is to examine whether the BRICs money demand functions have asymmetric adjustment toward equilibrium using threshold autoregressive (TAR, Enders and Granger, 1998) and momentum-threshold autoregressive (M-TAR, Enders and Granger, 1998; Enders and Siklos, 2001) models. The TAR and M-TAR models were formerly developed by Tong (1983), and these asymmetric error-correction models extend the original cointegration tests in the presence of asymmetric adjustment. As a recent studies employing nonlinear adjustment, Maki and Kitasaka (2006) investigate the long-run equilibrium relationship among money, income, prices, and interest rates in Japan by the threshold cointegration test, which allows for asymmetric adjustment, introduced by Enders and Siklos (2001). In the present paper, the threshold cointegration test introduced by Enders and Siklos (2001) is used, who expanded the Engle and Granger (1987) test into allowing for asymmetric adjustment toward equilibrium. Their proposed TAR model allows the degree of autoregressive decay to depend on the state of the variables.

Most models of the past empirical research addressing the issue of equilibrium have not taken into account the asymmetric properties of the adjustment process in money demand. Since Enders and Granger (1998) and Enders and Siklos (2001) proposed the asymmetrical TAR and M-TAR cointegration tests, discussing macroeconomic variables by applying nonlinear models are going to be the mainstream. Moreover, most developing countries of the demand for money have received less attention and separately investigated each country of the BRICs. Thus, this present empirical study contributes significantly to this field of research because, firstly, it determines whether stable money demand functions exist in the BRICs for which we use TAR and M-TAR cointegration tests. Second, this study is the first attempt to examine whether the BRICs money demand functions have asymmetric adjustment toward equilibrium. Third, we apply asymmetrical error-correction models to describe the money demand of the BRICs and the function could be served as the guideline for macro policy. To the best of our knowledge, our paper is the first attempt to investigate nonlinearity in the long-run cointegration relationship of the money demand function for BRICs.

METHODOLOGY

We specifically employ the threshold cointegration approach elaborated by Enders and Granger (1998), Enders and Siklos (2001). This is indeed a residual-based two-staged estimation as developed by Engle and Granger (1987). As an assumption of the tests for threshold cointegration, consider M2 series, denoted as y_t, and a set of n macroeconomic variables, $X_t = (x_{1t}, x_{2t}, ..., x_{nt})'$, however, in this study macroeconomic variables are real GDP, real exchange rates, interest rates. Suppose both y_t and x_t are I(1) series, and are linearly cointegrated with only one cointegrated relation, the long-run equilibrium relationship is given by:

$$y_t = \beta_0 + \beta_1 x_{1t} + \beta_2 x_{2t} + ... + \beta_n x_{nt} + \mu_t \quad (1)$$

where β_0 is a constant, $\beta_0,\ldots,\beta_n$ are estimated parameters and μ_t is the disturbance term that may be serially correlated. The existence of the long-run equilibrium relationship involves the stationarity of μ_t. To investigate the stationarity of μ_t, whether $-2 < \rho < 0$ has to be tested for in the second step procedure given by:

$$\Delta\mu_t = \rho\mu_{t-1} + \varepsilon_t \qquad (2)$$

where ε_t is a white-noise disturbance and the residuals from the regression model are used to estimate $\Delta\mu_t$. Rejection the null hypothesis of no cointegration (that is, accepting the alternative hypothesis $-2 < \rho < 0$) implies that the residuals in Equation 2 are stationary with mean zero. Hence, the long-run equilibrium relationship (Equation 1) with symmetric adjustment (Equation 2) is accepted.

The standard cointegration framework assuming symmetric adjustment toward equilibrium in Equation 2 is misspecified if the adjustment process is asymmetric. A formal way to introduce asymmetric adjustment is to let the deviation from the long-run equilibrium in Equation 1 behave as a TAR process. Enders and Siklos (2001) proposed test of threshold cointegration such that the residuals from Equation 1 are estimated in the form

$$\Delta\mu_t = I_t\rho_1\mu_{t-1} + (1 - I_t)\rho_2\mu_{t-1} + \varepsilon_t \qquad (3)$$

where I_t is the Heaviside indicator such that

$$I_t = \begin{cases} 1 & if \quad \mu_{t-1} \geq \tau \\ 0 & if \quad \mu_{t-1} < \tau \end{cases} \qquad (4)$$

and τ is the threshold value. If $\mu_{t-1} \geq \tau$, the value of $I_t = 1$, and the speed of adjustment in Equation 3 is ρ_1. Instead, if $\mu_{t-1} < \tau$, $I_t = 0$, the speed of adjustment is ρ_2. A necessary and sufficient condition for $\{\mu_t\}$ to be stationary is $-2 < (\rho_1, \rho_2) < 0$. The threshold parameter τ, which is restricted to the ranges of the remaining 70% of μ_t when the largest and smallest 15% values are discarded, is selected as an unknown value so as to minimize the sum of the squared residuals obtained from Equation 3 (Chan, 1993 for details). Enders and Granger (1998) also indicated that if the $\{\mu_t\}$ sequence is stationary, the least squares estimates of ρ_1 and ρ_2 have an asymptotic multivariate normal distribution if the value of the threshold is known (or consistently estimated). When the adjustment process (Equation 3) is serially correlated, Equation 3 is re-written as:

$$\Delta\mu_t = I_t\rho_1\mu_{t-1} + (1 - I_t)\rho_2\mu_{t-1} + \sum_{i=1}^{p}\gamma_i\Delta\mu_{t-i} + \varepsilon_t \qquad (5)$$

although it is possible that γ_i is asymmetric, for the sake of simplicity, this case is not considered, as in Enders and Siklos (2001).

Instead of estimating Equation 3 with the Heaviside indicator depending on the level of μ_{t-1}, the decay could also be allowed depending on the previous period's change in μ_{t-1}. In this case, the Heaviside indicator of Equation 4 becomes:

$$I_t = \begin{cases} 1 & if \quad \Delta\mu_{t-1} \geq \tau \\ 0 & if \quad \Delta\mu_{t-1} < \tau \end{cases} \qquad (6)$$

where τ is the threshold value. The Heaviside indicator could then be specified as $I_t = 1$ if $\Delta\mu_{t-1} \geq \tau$ and $I_t = 0$ if $\Delta\mu_{t-1} < \tau$. According to Enders and Granger (1998), this model is especially valuable when adjustment is asymmetric such that the series exhibits more "momentum" in one direction than the other. This model is termed M-TAR model. The TAR model is designed to capture asymmetrically "deep" movements in the series of the deviations from the long-run equilibrium, for example, if positive deviations are more prolonged than negative deviations. On the other hand, the M-TAR model is useful to capture the possibility of asymmetrically "steep" movements in the series. In the TAR model if $-1 < \rho_1 < \rho_2 < 0$, then the negative phase of μ_t will tend to be more persistent than the positive phase. For the M-TAR model, if for example $|\rho_1| < |\rho_2|$ the model exhibits little decay for positive $\Delta\mu_{t-1}$ but substantial decay for negative $\Delta\mu_{t-1}$. This means that increases tend to persist but decreases tend to revert quickly toward the attractor.

Finally, we can perform a number of statistical tests on the estimated coefficients in order to ascertain whether the variables are cointegrated and, in such a case, if the adjustment is symmetric or not. Enders and Siklos (2001) proposed two tests, called the Φ and *t*-max statistics. The Φ statistic using the *F*-statistic involves procedure testing for the null hypothesis $\rho_1 = \rho_2 = 0$, and *t*-max statistic using a *t*-statistic requires the test for the null hypothesis with the largest $\rho_i = 0$ between ρ_1 and ρ_2. If the null hypothesis with no cointegration is rejected, the null hypothesis $\rho_1 = \rho_2$ can be tested for with a standard *F*-statistic because the system is stationary. The equilibrium relationship with symmetric adjustment is accepted when the null hypothesis with no cointegration is rejected and the null hypothesis $\rho_1 = \rho_2$ is not rejected. In this case, the Engle–Granger (E-G) test for cointegration is a special case of Equation 3.

RESULTS

The data used in this study consist of quarterly observations on the natural logarithm of the real M2 money balance (lnM2), the natural logarithm of real gross domestic product (lnGDP), the natural logarithm of three months time deposit rate (lnTD), and the natural logarithm of the real exchange rate (lnEX) from the 1994 to 2008. Real money balance is obtained by deflating M2 by the consumer price index, and real gross domestic product is also obtained by deflating GDP by the consumer price index.

Table 1. The results of nonlinear unit root test - KSS test.

t- Statistic (Level)	lnM2[2]	lnGDP[3]	lnEX[4]	lnTD[5]
China	-1.5865	1.9267	-2.5058	-2.6120
India	1.0177	0.7954	-2.0416	-1.0414
Brazil	-0.5040	-0.9653	-1.7703	-2.5459
Russia	2.0123	-0.3323	-1.4626	-1.9801

1. The 0.01, 0.05, 0.1 asymptotic null critical values for KSS tests are -3.48, -2.93 and -2.66, respectively (Kapetanios et al., 2003).
2. lnM2 denotes the natural logarithm of the real M2 money balance.
3. lnGDP denotes the natural logarithm of real gross domestic product.
4. lnEX denotes the natural logarithm of the real exchange rate.
5. lnTD denotes the natural logarithm of three months time deposit rate.

The data were collected from the International Financial Statistics (IFS), AREMOS, and DATASTREAM databases.

Nonlinear unit root test

Recently, there is a growing consensus that macroeconomic variables might exhibit nonlinearities, and that conventional tests for stationarity, such as the ADF unit root test, have lower power in detecting the mean reverting (stationary) tendency of the series. For this reason, stationarity tests in a nonlinear framework must be applied.

This study employs the nonlinear stationary test advanced by Kapetanios, Shin, and Snell (2003, henceforth denoted as KSS test) to determine if the real M2 money balance (lnM2), real income (lnGDP), real exchange rate (lnEX), and deposit rate (lnTD) for the BRICs are nonlinear stationary. Table 1 presents the results of KSS (2003) nonlinear stationary test, which shows that all variables considered in this study are integrated of order one series, $I(1)$, at least at the 10% significant level.

The results indicate that the null of a unit root is not rejected against the nonlinear stationary alternatives for all variables.

Threshold cointegration tests

We found nonlinear relationship exist real M2 money balance (lnM2), real GDP (lnGDP), real exchange rate (lnEX), and deposit rate (lnTD) when we use KSS unit root test. Therefore, we go for threshold cointegration tests, Equation 1 was estimated using ordinary least squares (OLS) and saved the residuals in the sequence {μ_t}. For each type of asymmetry, we set the indicator function I_t according to Equation 4 or Equation 6 and estimated an equation in the form of Equation 5. The Akaike Information Criterion (AIC) and Schwartz Bayesian Information Criterion (SBC) were used to select the most appropriate lag length and to determine whether the adjustment mechanism is best captured as a TAR or M-TAR process. The results of the threshold cointegration test with zero and consistent estimate of the threshold are reported in Table 2.

For comparison purposes, the first rows of Table 2 present the E-G's cointegration test results. When we conduct the traditional linear E-G cointegration test, the null hypothesis can be rejected. In other words, there is one cointegration among all variables for each country. Notice that the AIC and SBC select the asymmetric models over the linear adjustment models for all countries, moreover, diagnostic checking of the residuals of the E-G's models show evidence of serial correlation. Thus, TAR and M-TAR models are more appropriate than the E-G's models. We also find that the consistent estimate of the threshold of TAR and M-TAR models with the AIC and SBC as the selection standards are superior to the TAR and M-TAR models with the threshold value of zero. As shown in Table 2, In the China case, based on AIC and SBC, the TAR model with the consistent estimate of the threshold is selected and the null hypothesis of $\rho_1 = \rho_2 = 0$ can be reject at the 10% significance level, whereas, the M-TAR model with the consistent estimate of the threshold are selected and the null hypotheses of $\rho_1 = \rho_2 = 0$ can be reject at the 1% significance level for India, Brazil, and Russia.

However, there is no reason to presume that the threshold is identically equal to zero. The consistent threshold estimates of 0.1365, 0.0481, 0.0497, and -0.2876 are obtained for China, India, Brazil, and Russia, respectively. We fail assuming linear adjustment or allowing for asymmetric adjustment using a threshold value of zero for the BRICs and find that there is a strong evidence of long-run money demand function for the BRICs. In addition, we test for symmetric versus asymmetric adjustment using the standard *F*-statistic. For India, Brazil, and Russia, the null hypothesis of symmetric adjustment is rejected at the 1% significance level. Besides, there is evidence that $|\rho_1| > |\rho_2|$ implying that the speed of adjustment is more rapid for positive than for negative discrepancies. For example, the real rate of the China

Table 2. The results of cointegration tests.

Country	Model	Lag	τ	ρ_1	ρ_2	AIC/SBC[1]	$\rho_1 = \rho_2 = 0$ [2]	$\rho_1 = \rho_2$ [3]	*Q*(4)[4]	Flag
China	E-G	5		-0.1954**		-65.2967/			98.0810	TAR
				(-2.1657)		-52.9683			[0.0000]	
	TAR	5	0	-0.3517***	-0.1370	-67.7576/	4.6689	1.7955	2.3551	
				(-2.9859)[5]	(-1.0778)[6]	-55.2683			[0.6708]	
			0.1365	-0.4147***	-0.1152	-70.0635/	5.9155*	3.8829	3.2969	
				(-3.4123)	(-1.0081)	-57.5742			[0.5094]	
	M-TAR	5	0	-0.2636**	-0.2382*	-65.7013/	3.6111	0.0242	2.2900	
				(-2.3240)	(-1.7194)	-53.2120			[0.6826]	
			-0.0846	-0.2518***	-0.8659*	-67.8981/	4.7431	1.9197	2.4082	
				(-2.7280)	(-1.9189)	-55.4088			[0.6611]	
India	E-G	2		-1.1309***		-94.9431/			10.1050	M-TAR
				(-5.1388)		-87.6285			[0.0390]	
	TAR	2	0	-0.9589***	-1.2373***	-98.1988/	14.4461***	1.3565	7.5147	
				(-3.5889)	(-5.3158)	-90.7982			[0.1111]	
			-0.0419	-0.9380***	-1.3299***	-99.5907/	15.5266***	2.6898	8.6851	
				(-3.7983)	(-5.4647)	-92.1901			[0.0695]	
	M-TAR	2	0	-0.9883***	-1.2622***	-98.0996/	14.3703***	1.2629	7.5985	
				(-3.8613)	(-5.1908)	-90.6990			[0.1074]	
			0.0481	0.0694	-1.0190***	-109.6553/	24.3684***	13.6004***	5.7468	
				(-0.1827)	(-5.2228)	-102.2547			[0.2189]	
Brazil	E-G	1		-0.4286**		-98.7379/			8.7616	M-TAR
				(-2.5639)		-94.9955			[0.0670]	
	TAR	1	0	-0.5844***	-0.2355	-98.7629/	4.3234	1.9391	7.6288	
				(-2.9256)	(-1.0908)	-93.1493			[0.1062]	
			0.0443	-0.6667***	-0.1750	-100.8533/	5.5174	4.0284	6.1906	
				(-3.3214)	(-0.8522)	-95.2397			[0.1854]	
	M-TAR	1	0	-0.5677***	-0.0547	-100.6774/	5.4149	3.8491	5.2355	
				(-3.2067)	(-0.2185)	-95.0638			[0.2640]	
			0.0497	-1.2243***	-0.2193	-111.6794/	12.6060***	16.4329***	3.5559	
				(-5.0211)	(-1.4276)	-106.0658			[0.4694]	
Russia	E-G	2		-0.7625*		137.9187/			12.6160	M-TAR
				(-2.8726)		145.2333			[0.0130]	
	TAR	2	0	-0.3144	-0.9107***	138.8633/	5.6618	1.7571	1.1959	
				(-0.7459)	(-3.3649)	146.2639			[0.8788]	
			-0.5089	-0.2616	-1.0343***	137.3828/	6.5310*	3.18934	1.8189	
				(-0.7007)	(-3.6120)	144.7834			[0.7690]	
	M-TAR	2	0	0.0489	-0.9875***	135.2674/	7.8215*	5.3158	3.8379	
				(-0.1147)	(-3.8417)	142.6680			[0.4284]	
			-0.2876	-0.0837	-4.1528***	125.6723/	28.4104***	16.4861***	1.358	
				(-1.0337)	(-23.6797)	133.0729			[0.8515]	

1. AIC=T*ln(RSS)+2*n ; and SBC=T*ln(RSS)+n*ln(T), where n = number of regressors and T = number of usable observations. RSS is the residual sum of squares. 2. This test follows a non-standard distribution so the test statistics are compared with critical values reported by Enders and Siklos (2001). 3. The numbers reported in this column are *F*-statistics of symmetric adjustment. The critical values are taken from Enders and Siklos (2001). 4. *Q*(4) is the Ljung-Box *Q*-statistic for the joint hypothesis of no serial correlation among the first residuals. 5. Entries in parentheses in this column are the *t*-statistics for the null hypothesis $\rho_1 = 0$ and $\rho_2 = 0$ Critical values are taken from Enders and Granger (1998). 6. Numbers in brackets are *p*-value. 7. The ***, **, and * indicate significance at the 0.01, 0.05 and 0.1 levels, respectively.

converges to its long-run equilibrium τ at the rate of 41.47% for a positive deviation and 11.52% for a negative deviation. It is reasonable to conclude that the money demand functions in the BRICs follow nonlinear adjustment and the adjustment mechanisms of India, Brazil, and Russia are asymmetric.

Error-correction models

Before examining a nonlinear error-correction model, we investigate the appropriateness of a linear ECM. Based on the estimates of the long-run demand for money, the SBC was used to select the short-run dynamic model and the order of autoregression. The linear ECM for money demand with AR(4) is specified as follows:

$$\Delta(\ln M2)_t = b_0 + b_1 ecm_{t-1} + \sum_{i=1}^{4} b_{1i}\Delta(\ln M2)_{t-i} + \sum_{i=0}^{4} b_{2i}\Delta(\ln GDP)_{t-i} + \sum_{i=0}^{4} b_{3i}\Delta(\ln EX)_{t-i} + \sum_{i=0}^{4} b_{4i}\Delta(\ln TD)_{t-i} + z_t \quad (7)$$

Where, $ecm_t = \ln M2_t - \hat{\beta}_0 - \hat{\beta}_1 \ln GDP_t - \hat{\beta}_2 \ln EX_t - \hat{\beta}_3 \ln TD_t$, it is the equilibrium error normalized on $\ln M2_t$, b_1 is the adjustment coefficient of the equilibrium error, which is expected to be negative. Not all of the coefficients in Equation 7 may be statistically significant in practice, and greater efficiency may actually be gained by removing the insignificant coefficients. We exclude those insignificant variables as long as their elimination does not produce evidence of serial correlation based on a *Q*-statistic at four lags. Using this procedure, variables are included, even when they are insignificant, if their deletion has resulted in serial correlation. The first and second columns of Table 3 to Table 6 report the estimation results of Equation 7. The linear ECM estimates appear to be reasonable with expected signs for India, Brazil, and Russia. The negative coefficient of the error-correction term reconfirms that the short-run adjustment moves the demand for money towards the long-run equilibrium.

However, Muscatelli and Spinelli (1996), Wolters et al. (1998) made the argument that the error correction to the short-run dynamics may not follow a linear process. First, OLS was used to estimate the long-run relationship is given by:

$$(\ln M2)_t = \beta_0 + \beta_1(\ln GDP)_t + \beta_2(\ln EX)_t + \beta_3(\ln TD)_t + \varepsilon_t \quad (8)$$

using Equation 8, in the case of China, the estimated asymmetric error-correction equations with the consistent estimates of the threshold is expressed as follows:

$$\Delta(\ln M2)_t = \alpha_0 + \gamma_1 S\_Plus_{t-1} + \gamma_2 S\_Minus_{t-1} + \sum_{i=1}^{k_1}\theta_{1i}\Delta(\ln M2)_{t-i} + \sum_{i=0}^{k_2}\theta_{2i}\Delta(\ln GDP)_{t-i} + \sum_{i=0}^{k_3}\theta_{3i}\Delta(\ln EX)_{t-i} + \sum_{i=0}^{k_4}\theta_{4i}\Delta(\ln TD)_{t-i} + \nu_t \quad (9)$$

Where, $S\_Plus_{t-1} = I_t(\ln M2_{t-1} - 1.2031 + 0.1821\ln GDP_{t-1} - 3.5264\ln EX_{t-1} + 0.1015\ln TD_{t-1})$

I_t = Heaviside indicator function, obtained by applying Chan's method to each country. The Heaviside indicator could be specified as $I_t = 1$ if $\varepsilon_{t-1} \geq 0.1365$ and $I_t = 0$ if $\varepsilon_{t-1} < 0.1365$. ν_t is a white-noise disturbance. We apply the SBC criterion to determine the appropriate lag lengths and empirically find that, for all cases, the four lag lengths of and are all four (that is, $k_1 = k_2 = k_3 = k_4 = 4$). The estimated asymmetric error-correction models with consistent estimate of thresholds are shown in the last two columns of Table 3 to Table 6. The estimated coefficients of S_Plus_{t-1} and S_Minus_{t-1} determine the speed of adjustment for positive and negative deviations from fundamental values, respectively.

For the adjustments towards long-run equilibrium in China, Table 3 shows that there are 2.16% (5.91%) adjustments to the equilibrium level when differences in the previous disequilibrium term are above (below) the threshold value of 0.1365 and the adjustments are symmetric. For India, Table 4 shows that there are 2.99% (7.56%) adjustments to the equilibrium level when differences in the previous disequilibrium term are above and below the threshold value of 0.0481 and the adjustments are asymmetric. Finally, Table 6 shows that there are 0.37% (428.18%) adjustments in the Russia to the equilibrium level when differences in the previous disequilibrium term are above (below) the threshold value of -0.2876 and the adjustments are asymmetric. These results indicate that negative deviations from fundamental values are eliminated quicker than positive deviations.

More specifically, the speeds of adjustment towards long-run equilibrium in Russia are much faster in the lower regime than in the higher regime. However, for Brazil, Table 5 shows that there are 53.45% (4.52%)

Table 3. Estimates of the error-correction models for China.

Variable	Symmetric[2]		Asymmetric[3]	
Constant	0.0214	(0.0054)***	0.0244	(0.0063)***
$\Delta lnM2_{t-3}$	0.6340	(0.1087)***	0.6119	(0.1034)***
$\Delta lnM2_{t-4}$	-0.2717	(0.1143)**	-0.2821	(0.1152)**
$\Delta lnEX_{t-2}$	-0.4762	(0.1614)***	-0.4846	(0.1452)***
$\Delta lnTD_{t-2}$	-0.0542	(0.0244)**	-0.0543	(0.0225)**
ECM_{t-1}	0.0409	(0.0170)**		
S_Plus_{t-1}			0.0216	(0.0259)
S_Minus_{t-1}			0.0591	(0.0260)**
Adj. R^2	0.5634		0.5669	
RSS[5]	0.0079		0.0078	
J-B[6]	4.3634	[0.1129]	2.5551	[0.2787]
$Q(4)$[7]	4.1119	[0.1280]	3.7076	[0.1566]
ARCH(4)[8]	3.7193	[0.4453]	3.2888	[0.5107]
Variance ratio[9]		0.9779		

Numbers in parentheses and brackets are standard errors and *p*-value, respectively.

2. Symmetric error-correction model:

$$\Delta(\ln M2)_t = a_0 + b_1 ecm_{t-1} + \sum_{i=1}^{4} b_{1i}\Delta(\ln M2)_{t-i} + \sum_{i=0}^{4} b_{2i}\Delta(\ln GDP)_{t-i} + \sum_{i=0}^{4} b_{3i}\Delta(\ln EX)_{t-i} + \sum_{i=0}^{4} b_{4i}\Delta(\ln TD)_{t-i} + z_t$$

3. Asymmetric error-correction model:

$$\Delta(\ln M2)_t = \alpha_0 + \gamma_1 S\_Plus_{t-1} + \gamma_2 S\_Minus_{t-1} + \sum_{i=1}^{k_1}\theta_{1i}\Delta(\ln M2)_{t-i} + \sum_{i=0}^{k_2}\theta_{2i}\Delta(\ln GDP)_{t-i} + \sum_{i=0}^{k_3}\theta_{3i}\Delta(\ln EX)_{t-i} + \sum_{i=0}^{k_4}\theta_{4i}\Delta(\ln TD)_{t-i} + \nu_t$$

where $S\_Plus_{t-1} = I_t\hat{\mu}_{t-1}$, $S\_Minus_{t-1} = (1-I_t)\hat{\mu}_{t-1}$ such that $I_t = 1$ if $\mu_{t-1} \geq 0.1365$, $I_t = 0$ if $\mu_{t-1} < 0.1365$ and ν_t is a white-noise disturbance.

4. The *** and ** indicate significance at the 0.01 and 0.05 levels, respectively.

5. RSS is sum of squared residuals.

6. J-B is the Jarque-Bera test of normality for the residual.

7. $Q(4)$ is the Ljung-Box autocorrelation tests for the residual.

8. ARCH(4) is the autoregressive conditional heteroscedasticity test of Engle (1982) and has χ^2 distribution with 4 degrees of freedom.

9. Variance ratio= σ_{NL}^2/σ_L^2, σ_{NL}^2 is the variance for the residual of asymmetric model, σ_L^2 is the variance for the residual of symmetric model.

adjustments to the equilibrium level when differences in the previous disequilibrium term are above and below the threshold value of 0.0497 and the adjustments are asymmetric. These indicate that positive deviations from fundamental values are eliminated quicker than negative deviations. Furthermore, we found that the coefficients on the error-correction terms are small except for that on S_Plus_{t-1} in the equation for Brazil and S_Minus_{t-1} in the equation for Russia. Other adjustments are small and statistically insignificant. In particular, there are only 0.37% adjustments in Russia to revert to the equilibrium level.

For comparison, estimates of both symmetric and asymmetric error-correction models are presented. Estimates for the asymmetric adjustments are presented in the last two columns, followed by the estimates from the symmetric error-correction models. For symmetric and asymmetric error- correction models, the Ljung-Box's Q-statistic fails to reject the hypothesis of no auto-correlation in residuals. In addition, the ARCH statistic of Engle (1982) fails to reject the hypothesis of no autoregressive conditional heteroscedasticity in residuals except for the linear ECM of Russia is 28.3764. The Jarque-Bera (J-B) statistic fails to reject the hypothesis of normality in residuals except for the linear ECM of Russia. However, the variance ratios are smaller than 1 for each country, we conjecture that a nonlinear model may be appropriate to describe the dynamics of money

Table 4. Estimates of the error-correction models for India.

Variable	Symmetric[2]		Asymmetric[3]	
Constant	0.0266	(0.0036)***	0.0249	(0.0040)***
$\Delta lnM2_{t-4}$	-0.2795	(0.1277)**	-0.2447	(0.1324)*
$\Delta lnGDP_{t-2}$	0.1595	(0.0440)***	0.1777	(0.0477)***
$\Delta lnTD_{t-2}$	-0.0740	(0.0243)***	-0.0733	(0.0243)***
ECM_{t-1}	-0.0527	(-0.0452)		
S_Plus_{t-1}			0.0299	(-0.0946)
S_Minus_{t-1}			-0.0756	(-0.0508)
Adj. *R*-squared	0.4138		0.4139	
RSS[5]	0.0090		0.0088	
J-B[6]	4.4842	[0.1062]	3.6867	[0.1583]
Q(4)[7]	1.1621	[0.8843]	1.8236	[0.7682]
ARCH(4)[8]	7.2604	[0.1228]	6.5780	[0.1599]
Variance ratio[9]		0.9755		

1. Numbers in parentheses and brackets are standard errors and *p*-value, respectively.
2. Symmetric error-correction model:

$$\Delta(\ln M2)_t = a_0 + b_1 ecm_{t-1} + \sum_{i=1}^{4} b_{1i}\Delta(\ln M2)_{t-i} + \sum_{i=0}^{4} b_{2i}\Delta(\ln GDP)_{t-i} + \sum_{i=0}^{4} b_{3i}\Delta(\ln EX)_{t-i} + \sum_{i=0}^{4} b_{4i}\Delta(\ln TD)_{t-i} + z_t$$

3. Asymmetric error-correction model:

$$\Delta(\ln M2)_t = \alpha_0 + \gamma_1 S\_Plus_{t-1} + \gamma_2 S\_Minus_{t-1} + \sum_{i=1}^{k_1} \theta_{1i}\Delta(\ln M2)_{t-i} + \sum_{i=0}^{k_2} \theta_{2i}\Delta(\ln GDP)_{t-i} + \sum_{i=0}^{k_3} \theta_{3i}\Delta(\ln EX)_{t-i} + \sum_{i=0}^{k_4} \theta_{4i}\Delta(\ln TD)_{t-i} + \nu_t$$

where $S\_Plus_{t-1} = I_t\hat{\mu}_{t-1}$, $S\_Minus_{t-1} = (1-I_t)\hat{\mu}_{t-1}$ such that $I_t = 1$ if $\Delta\mu_{t-1} \geq 0.0481$, $I_t = 0$ if $\Delta\mu_{t-1} < 0.0481$ and ν_t is a white-noise disturbance.

4. The ***, **, and * indicate significance at the 0.01, 0.05 and 0.1 levels, respectively.
5. RSS is sum of squared residuals.
6. J-B is the Jarque-Bera test of normality for the residual.
7. *Q*(4) is the Ljung-Box autocorrelation tests for the residual.
8. ARCH (4) is the autoregressive conditional heteroscedasticity test of Engle (1982) and has χ^2 distribution with 4 degrees of freedom.
9. Variance ratio= $\sigma_{NL}^2 / \sigma_L^2$, σ_{NL}^2 is the variance for the residual of asymmetric model, σ_L^2 is the variance for the residual of symmetric model.

money demand. More specifically, the variance ratio of Russia is 0.0103. Because the variance ratio of Russia is much smaller than 1, we argue that the linear ECM of Russia must be misspecification.

These empirical Findings from Table 3 to Table 6 indicate that the asymmetric ECM may be appropriate to describe the dynamics of the BRICs money demand. In other words, the short-run dynamics towards the long-run equilibrium of the money demand in the BRICs follow nonlinear adjustment. It implies that, within the context of money demand, households and the government may respond differently when the economy is in a different regime under the specific threshold and can provide useful information about portfolio allocation. Although the linear ECM specification gives similar estimation on parameters, the nonlinear specification outperforms the linear ECM when the country experiences a volatile economic condition as BRICs. Our result, the estimated coefficient of error correction term, indicates that there is cointegration among variables in money demand function. The results also reveal that the estimated elasticity coefficients of real income are positive and negative as expected. For exchange rate, we obtain the results of negative coefficients which support the currency substitution symptom in BRICs. It is worth noting that, in long-run, even the coefficient of exchange rate has negative sign, supporting the currency substitution phenomenon in BRICs.

Table 5. Estimates of the error-correction models for Brazil.

Variable	Symmetric[2]		Asymmetric[3]	
Constant	0.0175	(0.0050)***	0.0188	(0.0048)***
$\Delta lnM2_{t-1}$	0.2754	(0.1499)*	0.4138	(0.1536)***
$\Delta lnM2_{t-2}$	0.2947	(0.1628)*	0.2976	(0.1537)*
$\Delta lnM2_{t-4}$	-0.3863	(0.1265)***	-0.3866	(0.1194)***
$\Delta lnGDP_{t-1}$	0.3765	(0.1940)*	0.4095	(0.1837)**
$\Delta lnGDP_{t-2}$	0.4230	(0.2293)*	0.4240	(0.2165)*
$\Delta lnEX_{t-2}$	-0.1215	(0.0453)**	-0.1142	(0.0428)**
$\Delta lnTD_{t-2}$	0.0360	(0.0194)*	0.0446	(0.0186)**
ECM_{t-1}	-0.0014	(-0.0939)		
S_Plus_{t-1}			-0.5345	(0.2460)**
S_Minus_{t-1}			0.0452	(-0.0909)
Adj. R^2	0.3111		0.3861	
RSS[5]	0.0203		0.0176	
J-B[6]	0.1429	[0.9311]	0.9349	[0.6266]
$Q(4)$[7]	2.0746	[0.7220]	2.8011	[0.5916]
ARCH(4)[8]	5.2508	[0.2625]	3.2870	[0.5110]
Variance ratio[9]		0.8658		

1. Numbers in parentheses and brackets are standard errors and *p*-value, respectively.
2. Symmetric error-correction model:

$$\Delta(\ln M2)_t = a_0 + b_1 ecm_{t-1} + \sum_{i=1}^{4} b_{1i}\Delta(\ln M2)_{t-i} + \sum_{i=0}^{4} b_{2i}\Delta(\ln GDP)_{t-i} + \sum_{i=0}^{4} b_{3i}\Delta(\ln EX)_{t-i} + \sum_{i=0}^{4} b_{4i}\Delta(\ln TD)_{t-i} + z_t$$

3. Asymmetric error-correction model:

$$\Delta(\ln M2)_t = \alpha_0 + \gamma_1 S\_Plus_{t-1} + \gamma_2 S\_Minus_{t-1} + \sum_{i=1}^{k_1} \theta_{1i}\Delta(\ln M2)_{t-i} + \sum_{i=0}^{k_2} \theta_{2i}\Delta(\ln GDP)_{t-i} + \sum_{i=0}^{k_3} \theta_{3i}\Delta(\ln EX)_{t-i} + \sum_{i=0}^{k_4} \theta_{4i}\Delta(\ln TD)_{t-i} + \nu_t$$

where $S\_Plus_{t-1} = I_t\hat{\mu}_{t-1}$, $S\_Minus_{t-1} = (1-I_t)\hat{\mu}_{t-1}$ such that $I_t = 1$ if $\Delta\mu_{t-1} \geq 0.0497$, $I_t = 0$ if $\Delta\mu_{t-1} < 0.0497$ and ν_t is a white-noise disturbance.

4. The ***, **, and * indicate significance at the 0.01, 0.05 and 0.1 levels, respectively.
5. RSS is sum of squared residuals.
6. J-B is the Jarque-Bera test of normality for the residual.
7. $Q(4)$ is the Ljung-Box autocorrelation tests for the residual.
8. ARCH (4) is the autoregressive conditional heteroscedasticity test of Engle (1982) and has χ^2 distribution with 4 degrees of freedom.
9. Variance ratio= σ_{NL}^2/σ_L^2 , σ_{NL}^2 is the variance for the residual of asymmetric model, σ_L^2 is the variance for the residual of symmetric model.

Conclusions

A significant volume of empirical studies has shown that financial sequence data are commonly subject to non-linear dynamic adjustments. Enders and Granger (1998) determined that when there is asymmetric adjustment to economic variables, traditional linear unit root tests and cointegration tests have low power. The information on the structure of money demand function is very important for policy makers in designing effective monetary policy. This study aims to empirically investigate the long-run equilibrium relationship among real M2 money balance, real income, real exchange rate and deposit rate in the BRICs using the asymmetrical TAR and M-TAR cointegration tests developed by Enders and Granger (1998) and Enders and Siklos (2001). A modified money demand function, motivated by the literature of currency substitution, is applied in our empirical analysis in which a real exchange rate variable is included in the function. The TAR and M-TAR cointegration methods provided strong evidence of the long-run equilibrium relationship characterized by asymmetric adjustment. Using data from 1994

Table 6. Estimates of the error-correction models for Russia.

Variable	Symmetric[2]		Asymmetric[3]	
constant	-0.0423	(0.0870)	0.0321	(0.0091)***
$\Delta \ln M2_{t-1}$	0.7197	(0.2536)***	-0.0106	(0.0290)
$\Delta \ln M2_{t-2}$	5.6747	(1.3888)***	-0.7519	(0.1809)***
$\Delta \ln GDP_{t-1}$	1.0888	(1.8634)	0.5963	(0.1940)***
$\Delta \ln GDP_{t-2}$	4.7176	(1.9152)**	0.4641	(0.2165)**
$\Delta \ln GDP_{t-3}$	0.5584	(1.9826)	0.4949	(0.2041)**
$\Delta \ln GDP_{t-4}$	4.4700	(1.8050)**	0.1145	(0.2020)
$\Delta \ln EX_{t-2}$	-2.1869	(0.8831)**	-0.0577	(0.0988)
ECM_{t-1}	-0.7645	(0.2635)***		
S_Plus_{t-1}			-0.0037	(0.0301)
S_Minus_{t-1}			-4.2818	(0.0665)***
Adj. R^2	0.3727		0.9934	
RSS[5]	9.9761		0.1028	
J-B[6]	207.4273	[0.0000]	0.6595	[0.7191]
$Q(4)$[7]	0.7361	[0.9468]	3.8722	[0.4236]
ARCH(4)[8]	28.3764	[0.0000]	0.6604	[0.9561]
Variance ratio[9]		0.0103		

1. Numbers in parentheses and brackets are standard errors and *p*-value, respectively.
2. Symmetric error-correction model:

$$\Delta(\ln M2)_t = a_0 + b_1 ecm_{t-1} + \sum_{i=1}^{4} b_{1i}\Delta(\ln M2)_{t-i} + \sum_{i=0}^{4} b_{2i}\Delta(\ln GDP)_{t-i} + \sum_{i=0}^{4} b_{3i}\Delta(\ln EX)_{t-i} + \sum_{i=0}^{4} b_{4i}\Delta(\ln TD)_{t-i} + z_t$$

3. Asymmetric error-correction model:

$$\Delta(\ln M2)_t = \alpha_0 + \gamma_1 S\_Plus_{t-1} + \gamma_2 S\_Minus_{t-1} + \sum_{i=1}^{k_1} \theta_{1i}\Delta(\ln M2)_{t-i} + \sum_{i=0}^{k_2} \theta_{2i}\Delta(\ln GDP)_{t-i} + \sum_{i=0}^{k_3} \theta_{3i}\Delta(\ln EX)_{t-i} + \sum_{i=0}^{k_4} \theta_{4i}\Delta(\ln TD)_{t-i} + \nu_t$$

where $S\_Plus_{t-1} = I_t\hat{\mu}_{t-1}$, $S\_Minus_{t-1} = (1-I_t)\hat{\mu}_{t-1}$ such that $I_t = 1$ if $\Delta\mu_{t-1} \geq -0.2876$, $I_t = 0$ if $\Delta\mu_{t-1} < -0.2876$ and ν_t is a white-noise disturbance.

4. The ***, **, and * indicate significance at the 0.01, 0.05 and 0.1 levels, respectively.
5. RSS is sum of squared residuals.
6. J-B is the Jarque-Bera test of normality for the residual.
7. $Q(4)$ is the Ljung-Box autocorrelation tests for the residual.
8. ARCH (4) is the autoregressive conditional heteroscedasticity test of Engle (1982) and has χ^2 distribution with 4 degrees of freedom.
9. Variance ratio= σ_{NL}^2/σ_L^2 , σ_{NL}^2 is the variance for the residual of asymmetric model, σ_L^2 is the variance for the residual of symmetric model.

1994 to 2007, this study found that the short-run dynamics towards the long-run equilibrium of the money demand in the BRICs follow nonlinear adjustment. Although the linear ECM specification gives similar estimation on parameters, the nonlinear specification out performs the linear ECM when judged by such diagnostic tests as serial correlation, the ARCH effect, variance ratio, and adjusted R^2. These findings offer a new piece of evidence supporting the existence of the long-run equilibrium relationship of the BRICs money demand function with asymmetric adjustment.

The estimated model in this study can provide useful policy guidelines to the BRICs' central banks in their quest for price stability and narrowing the divergence between potential output and actual output. It is argued that any persistent disequilibrium in the money market can bring about rising future prices and widening gap between actual and potential output. Thus, if the objectives of these countries are to minimize the output gap and price instability, they should avoid creating unnecessary disequilibrium in the money market. That is why the stable long-run relationship between the real demand for money and other macroeconomic variables serves as the guideline for macro policy. Therefore, this study is the first attempt to model symmetric and asymmetric error-correction models for the demand for money in the BRICs.

REFERENCES

Anderson RG, Rasche RH (2001). The Remarkable Stability of Monetary Base Velocity in the United States 1919-1999, Federal Reserve Bank of St. Louis Working Paper.

Arize AC (1994). Has M2 Demand in the USA Become Unstable? Evidence from an Error-Correction Model. Appl. Econ., 26: 957-967.

Austin D, Ward B, Dalziel P (2007). The Demand for Money in China 1987–2004: A Nonlinear Modeling Approach, China Econ. Rev., 18: 190-204.

Baba Y, Hendry DF, Starr, RM (1992). The Demand for M1 in the USA 1960–1988. Rev. Econ. Stud., 59: 25-61.

Bahmani-Oskooee M, Pourheydarian M (1990). Exchange Rate Sensitivity of Demand for Money and Effectiveness of Fiscal and Monetary Policies, Appl. Econ., 22: 917-925.

Ball L (2001). Another Look at Long-Run Money Demand, J. Monetary Econ., 47: 31-44.

Chan KS (1993). Consistency and Limiting Distribution of the Least Squares Estimator of a Threshold Autoregressive Model, Annals Stat., 21: 520-533.

Deng S, Liu B (1999). Modelling and Forecasting the Money Demand in China: Cointegration and Nonlinear Analysis. Annals Oper. Res., 87: 177-189.

Enders W, Granger CWF (1998). Unit Root Tests and Asymmetric Adjustment with an Example Using the Term Structure of Interest Rates. J. Bus. Econ. Stat., 16: 304-311.

Enders W, Siklos PL (2001). Cointegration and Threshold Adjustment, J. Bus. Econ. Stat., 19: 166-176.

Engle R (1982). Autoregressive Conditional Heteroscedasticity with Estimates of the Variance of United Kingdom Inflation. Econometrics, 50: 987-1007.

Engle RF, Granger CWJ (1987). Cointegration and Error Correction: Representation, Estimation and Testing. Econometrics, 55: 251-276.

Ericsson NR, Hendry DF, Prestwich KM (1998). The Demand for Broad Money in the United Kingdom 1878-1993, Scandinavian J. Econ., 100: 289-324.

Friedman BM, Kuttner KN (1992). Money, Income, Prices, and Interest Rates, Am. Econ. Rev., 82: 472-492.

Friedman M, Schwartz AJ (1991). Alternative Approaches to Analyzing Economic Data, Am. Econ. Rev., 81: 39-49.

Hafer R W, Jansen DW (1991). The Demand for Money in the United States: Evidence from Cointegration Tests, J. Money, Credit Banking 23: 155-168.

Harrison B, Vymyatnina Y (2005). Demand for Money During Transition: The Case of Russia, Working Paper, European University at Saint-Petersburg.

Johansen S (1988). Statistical Analysis of Cointegration Vectors, J. Econ. Dynamics Control, 12: 231-254.

Kapetanios G, Shin Y, Snell A (2003). Testing for a Unit Root in the Nonlinear STAR Framework, J. Econom., 112: 359-379.

MacDonald R, Taylor MP (1992). A Stable US Money Demand Function 1874-1975, Econ. Lett., 39: 191-198.

McNown R, Wallace MS (1992). Cointegration Tests of a Long-Run Relation Between Money Demand and the Effective Exchange Rate, J. Int. Money Financ., 11:107–114.

Miyao R (1996). Does a Cointegrating M2 Demand Relation Really Exist in the United States? J. Money, Credit, Banking, 28: 365–380.

Muscatelli VA, Spinelli F (1996). Modeling Monetary Trends in Italy Using Historical Data: the Demand for Broad Money 1861–1990, Econ. Inquiry, 34: 579-596.

Rao BB, Singh R (2006). Demand for Money in India: 1953-2003, Appl. Econ., 38: 1319-1326.

Stock JH, Watson MW (1993). A Simple Estimator of Cointegrating Vectors in Higher Order Integrated Systems, Econometric, 61: 783-820.

Tong H (1983). Threshold Models in Nonlinear Time Series Analysis, Springer-Verlag, New York.

Wilson D, Purushothaman R (2003). Dreaming With BRICs: The Path to 2050. Global Econ. Paper, 99: 1-24.

Wolters LT, Teräsvirta T, Lutkepohl H (1998). Modelling the Demand for M3 in the United Germany, Rev. Econ. Stat., 80: 399-409.