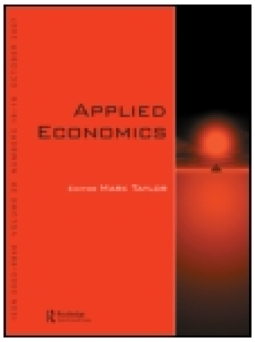

Applied Economics

ISSN: 0003-6846 (Print) 1466-4283 (Online) Journal homepage: http://www.tandfonline.com/loi/raec20

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To cite this article: B. Bhaskara Rao & Saten Kumar (2009) Cointegration, structural breaks and the demand for money in Bangladesh, Applied Economics, 41:10, 1277-1283, DOI: 10.1080/00036840701367671

To link to this article: https://doi.org/10.1080/00036840701367671

Published online: 11 Apr 2011.

Applied Economics, 2009, **41**, 1277–1283

Routledge
Taylor & Francis Group

Cointegration, structural breaks and the demand for money in Bangladesh

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This article allows for endogenous structural breaks in the cointegration equation and investigates if there is a stable demand for money for Bangladesh. We have used the Gregory and Hansen framework and found that there was an intercept shift and a well-determined and stable demand for money in Bangladesh exists.

I. Introduction

This article has three objectives viz., (1) to show the usefulness of some recent developments in the cointegration techniques which accommodate endogenous structural breaks in the underlying relationships (2) to illustrate this technique by estimating the demand for money for Bangladesh and by investigating if a long run demand for money relationship, in the presence of structural breaks, exists for Bangladesh and finally (3) to examine whether the money demand function for Bangladesh has become unstable due to financial deregulation and reforms of 1980s.[1]

Our first objective is important in that there is a persistent confusion between testing for unit roots in a variable and cointegration among a set of unit root variables with structural breaks. Although the test procedures are similar, conceptually they have different purposes. The third objective is also important because stability of the demand for money has implications for the choice of monetary policy instruments. According to Poole (1970) policy makers should target the rate of interest if the LM curve is unstable and target money supply if the IS curve is unstable. Since instability in LM is largely caused by instability in the money demand function, it is important to test for the stability of demand for money.

Compared to a vast literature on the demand for money for many countries, studies on demand for money in Bangladesh are limited. Furthermore, estimates of the demand for money that allow endogenous structural breaks are also limited for all countries. In this article, we shall use the Gregory and Hansen (1996a and 1996b) techniques that investigate structural breaks in the cointegrating relationships. Our estimates with this technique show that there is a stable cointegrating relationship between real narrow money, real income and nominal rate of interest in Bangladesh from 1980 to 2003. However, there was an intercept shift in this relationship, most probably in 1989. An important implication of our finding is that the Central Bank of Bangladesh should target money supply, instead of the rate of interest, as its instrument of monetary policy.

The outline of this article is as follows. Section II reviews some previous empirical studies on demand for money in Bangladesh. In Section III the Gregory and Hansen(GH) technique is explained and used for

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[1] We could have selected any relationship and data from any country to illustrate our technique. However, we have selected the demand for money in Bangladesh because relatively there are only a small number of empirical works on this topic.

Applied Economics ISSN 0003–6846 print/ISSN 1466–4283 online © 2009 Taylor & Francis
http://www.informaworld.com
DOI: 10.1080/00036840701367671

estimating cointegrating equations with endogenous structural breaks. Section IV presents empirical results and the summary and conclusions are in Section V. A limitation of this study is that it is not a comprehensive and all encompassing study on the demand for money in Bangladesh, Nevertheless, our specification and estimates are comparable to a very comprehensive recent study on the demand for money of a number of countries by Bahmani-Oskooee and Rehman (2002).

II. Empirical Studies on Bangladesh

There are only a handful of empirical studies on the demand for money for Bangladesh. Hossain (2006) recently estimated demand for narrow and broad money for Bangladesh using a totally outdated partial adjustment method (*PAM*) for the period 1973 to 2003. Siddiki (2000) used annual data from 1975 to 1995 to estimate the demand for real broad money (*M2*) with the bounds test approach, which was popularized to estimate demand for money functions by Bahmani-Oskooee and Rehman (2005). Ahmed (2001) studied the existence of a long-run demand for narrow and broad money functions for the period 1974 to 1995. Although these are pioneering studies for Bangladesh, each of these studies has limitations. Furthermore, in none of these studies the possibility of a structural break in the long-run cointegrating relationship, as in many other developing countries, has been investigated. Therefore, only for the sake of completeness, we shall briefly review these three works.

Hossain (2006) has ignored the implications of unit roots in the variables and used a totally outdated *PAM* framework to estimate the demand for money for 1973 to 2003 and sub-sample periods of 1977 to 2003, 1983 to 2003 and 1985 to 2003. His long run income elasticity estimates range from 1.14 for the entire sample period to 0.87 in the financial reform period of 1985 to 2003. Estimates of semi-interest rate elasticities are correctly and negatively signed and range from −0.13 in the whole sample period to −0.76 in 1983 to 2003. In the financial reforms period of 1985 to 2003, interest rate elasticity was −0.65. Although these estimates seem plausible and statistically significant, it is well known that his estimated *t*-ratios and other summary measures are over-estimated and unreliable.[2] Furthermore, the inappropriateness of using *PAM* dynamic adjustment was clearly highlighted by Taylor (1994). Another study by Hossain (1993) on the demand for money for Bangladesh contains similar drawbacks because he has used *PAM* to model the dynamics and ignored the unit roots in the variables. He has used in this study quarterly data from 1976Q1 to 89Q4 and found that the income elasticity for narrow money was low at 0.63.

Siddiki (2000) used annual data from 1975 to 1995 to estimate the demand for broad money (*M2*) for Bangladesh using the bounds test approach. His long run model corresponding to his the ARDL (2,0,2,0) formulation for the real per capita demand for broad money is[3]:

$$M = \underset{(7.73)^{**}}{-21.47} + \underset{(10.86)^{**}}{3.26}\, g + \underset{(4.50)^{**}}{0.088}\, r^d - \underset{(1.54)}{0.145}\, r^f \quad (1)$$

where M is the logarithm of real per capita broad money, g is the logarithm of real per capita income, r^d is domestic interest rates proxied by bank discount rate and r^f is the foreign interest rate, proxied by the unofficial exchange rate premiums as a percentage of unofficial exchange rates. *t*-ratios are below the coefficients.

However, Siddiki's estimate of income elasticity at 3.26 is high and seem to be implausible. It is expected that income elasticity to be around unity in the developing countries; see Sriram (1999). The implied interest rate elasticity has the expected negative sign and its magnitude is plausible. But the coefficient of the proxy for the effects of the foreign interest rate is insignificant at the conventional levels.

Ahmed (2002) estimated long run demand for narrow (M1) and broad money (M2) for the period 1974 to 1995. He has used the *PAM* adjustment framework and therefore has the same limitations of the study by Hossain (2006). His explanatory variables are per capita real income, real rate of interest, rate of inflation, degree of monetization and the real exchange rate. Inclusion of the real rate of interest gives the impression that the author wrongly mistook that the rate of interest should be real because the income variable is measured in real terms. His long run estimates of income elsaticities for M1 and M2 are, respectively, 0.8 and 1.2. The semi-interest rate elsaticities, respectively, are −0.04 and −0.003. However, since Ahmed measures the rate of interest in real terms it is difficult to take these estimates without reservations.

[2] The estimated adjusted $\bar{R}^2$ are all close to unity and the author did not report any measures to test for autocorrelation in the residuals.

[3] Significance at 1% is indicated by **.

Our brief review of these studies indicates is perhaps the only study that is econometrically satisfactory is that by Siddiki. However, his estimate of income elasticity at more than three is highly implausible. The other two studies by Hossain (2006) and Ahmed (2002) are econometrically unsatisfactory because they have ignored unit roots in the variables and their summary statistics are biased. Therefore, in what follows, we start with a clean slate and estimate the demand for narrow money in Bangladesh.

III. GH Methodology

At the outset of this section it may be noted that in none of the earlier studies on the demand for money for Bangladesh the time series variables were tested for unit roots.[4] We shall test the variables for unit roots later in this section and first explain the Gregory-Hansen procedure of testing for cointegration with endogenous structural breaks. Our specification of demand for money is simple and standard in which the demand for money (M1) is assumed to depend on income and the rate of interest. We ignore the foreign rates of interest because holding money in foreign exchange is not a realistic option to many in the developing countries. Our specification of demand for money is:

$$\ln M_t = \mu + a_1 \ln Y_t - a_2 r_t + e_t \tag{2}$$

where M is real narrow money, Y is real GDP, r is the nominal rate of interest and e is the error term.

The GH approach is an extension of similar tests for unit root tests with structural breaks, for example, by Zivot and Andrews (1992). Gregory and Hansen propose the cointegration tests which accommodates a single endogenous break in an underlying cointegrating relationship. The four models of Gregory and Hansen (1996a and 1996b) with assumptions about structural breaks and their specifications with two variables, for simplicity, are as follows:

Model 1: Level shift

$$Y_t = \mu_1 + \mu_2 f_{tk} + a_1 X_t + e_t \tag{3}$$

Model 2: Level shift with trend

$$Y_t = \mu_1 + \mu_2 f_{tk} + \beta_1 t + a_1 X_t + e_t \tag{4}$$

Model 3: Regime shift where intercept and slope coefficients change

$$Y_t = \mu_1 + \mu_2 f_{tk} + \beta_1 t + a_1 X_t + a_2 X_t f_{tk} + e_t \tag{5}$$

Model 4: Regime shift where intercept, slope coefficients and trend change

$$Y_t = \mu_1 + \mu_2 f_{tk} + \beta_1 t + \beta_2 t f_{tk} + a_1 X_t + a_2 X_t f_{tk} + e_t \tag{6}$$

where Y is the dependent and X is the independent variable, t is time subscript, ε is an error term, k is the break date and ϕ is a dummy variable such that:

$$f_{tk} = 0 \quad \text{if} \quad t \le k \text{ and } f_{tk} = 1 \text{ if } t > k \tag{7}$$

The null hypothesis of no cointegration with structural breaks is tested against the alternative of cointegration by the GH approach. The single break date in these models is endogenously determined. In all the previous studies on demand for money in Bangladesh, and in fact in many other countries, an important issue that was not addressed is that the cointegration relationship may have a structural break during the sample period. Therefore, we explore the stability of the demand for money with the Gregory–Hansen techniques. The GH demand for money specifications for the aforesaid four models, with structural breaks, are as follows:z

GH-I: Level shift

$$\ln M_t = \mu_1 + \mu_2 f_{tk} + a_1 \ln Y_t - a_2 r_t + e_t \tag{8}$$

GH-II: Level shift with trend

$$\ln M_t = \mu_1 + \mu_2 f_{tk} + \beta_1 t + a_1 \ln Y_t - a_2 r_t + e_t \tag{9}$$

GH-III: Regime shift where intercept and slope coefficients change

$$Y_t = \mu_1 + \mu_2 f_{tk} + \beta_1 t + a_1 \ln Y_t + a_{11} \ln Y_t f_{tk} - a_2 r_t - a_{22} r_t f_{tk} + e_t \tag{10}$$

GH-IV: Regime shift where intercept, slope coefficients and trend change

$$Y_t = \mu_1 + \mu_2 f_{tk} + \beta_1 t + \beta_2 t f_{tk} + a_1 \ln Y_t + a_{11} \ln Y_t f_{tk} - a_2 r_t - a_{22} r_t f_{tk} + e_t \tag{11}$$

The break date is found by estimating the cointegration equations for all possible break dates in the sample. We select a break date where the test statistic is the minimum or in other words the absolute ADF test statistic is at its maximum. GH have tabulated the critical values by modifying the MacKinnon (1991) procedure for testing cointegration in the Engle–Granger method for unknown breaks.

[4] The bounds test used by Siddiki does not require pre-testing the variables for unit roots.

IV. Empirical Results

We first tested for the presence of unit roots in our variables. The Augmented Dicky–Fuller test (ADF) is used for testing for the order of the variables. The time trend is included because it is significant in the levels and first differences of the variables. The computed test statistics for the levels and first differences of the variables are given in Table 1 below:

The null hypothesis of unit root cannot be rejected at the 5% level for the level variables of ln M, ln Y and r, but the null that their first differences have unit roots is clearly rejected. It is well known that the ADF test has a low power against the null. Therefore, since our ADF tests clearly indicate that the variables in their first differences are stationary (i.e. the null of unit roots is rejected) there is no point in wasting space by conducting alternative tests that have more power against the null. The definitions of variables and sources of data are in the appendix.

Table 1. ADF test for unit roots: levels and first difference of variables with intercept and linear trend

Variable	L	t-statistic	95% CV
ln M	0	−1.647	−3.594
? ln M	0	−4.097*	−3.603
ln Y	3	−2.263	−3.594
? ln Y	0	−6.869*	−3.603
r	4	−2.049	−3.594
? r	1	−3.730*	−3.603

Notes: L is the lag length of the first differences of the variables.
* indicates significance at 5% level. The sample period is 1973 to 2003.

The results for GH cointegration tests are given below in Table 2.

Table 2. Tests for cointegration with structural breaks 1973–2003

	Brake date	GH test statistic	5% Critical value	Reject H_0 of no cointegration?
GH-I	1989	−6.23601	−4.92	YES
GH-II	1988	−6.10633	−5.29	YES
GH-III	1989	−6.34941	−5.50	YES
GH-IV	1986	−6.59181	−6.00	YES

These results in Table 2 imply that in all the four models with structural breaks, there is cointegration between real narrow money, real income and the nominal rate of interest in Bangladesh. The brake date is 1989 in GH-I and GH-III, but different at 1988 and 1986 in GH-II and GH-IV, respectively. The null hypothesis of no cointegration is rejected in all the four models.

To select the best possible model we proceed to estimate the cointegrating equations for these four models with the Engle–Granger method. The first stage OLS equations are given below in Table 3. The estimates of these four models seem to imply that GH-I is the most plausible model for the following reasons. In GH-I, all the estimated coefficients are significant with the expected signs and magnitudes. The income elasticity of demand for money is 1.26 and the Wald-test could not reject the null that it is unity at the 5% level. The Wald-test computed ?^{2}(1) test statistic with p-value in the parenthesis is 2.237 (0.135) is insignificant.

In GH-II, the estimate of income elasticity has incorrect sign and insignificant at the conventional levels. In GH-III, the two income elasticities are implausible as one is very low (about 0.27) and the other a bit high (about 1.45) and the two interest rate coefficients are insignificant. Similarly in GH-IV, the income elasticity, after break, is insignificant and very high (about 5) while the other has incorrect sign.

Table 3. Cointegrating equations 1974–2003

	GH-I (DUM1989)		GH-II (DUM1988)		GH-III (DUM1989)		GH-IV (DUM1986)	
Intercept	1.914	(2.93)*	12.648	(2.86)*	5.144	(4.17)*	17.214	(3.17)*
Dum × Intercept	−0.368	(2.67)*	12.156	(2.79)*	0.771	(0.62)	−13.294	(1.02)
Trend			0.133	(2.40)*			0.183	(2.32)*
Dum × Trend							−0.205	(1.17)
ln Y_t	1.261	(7.23)*	−1.686	(1.40)	0.268	(0.73)	−2.963	(2.02)**
Dum × ln Y_t					1.449	(6.00)*	5.513	(1.48)
r_t	−0.030	(1.88)**	−0.035	(2.30)*	0.049	(1.61)	−0.019	(0.43)
Dum × r_t					−0.043	(1.58)	−0.031	(1.05)

Notes: Absolute t-ratios are in parentheses below the coefficients. Significance at 5% and 10% levels, respectively, is indicated with * and **. The year relevant for the dummy variable is indicated in the first row in the parentheses. DUM1989 means that the dummy is unity after that year and so on.

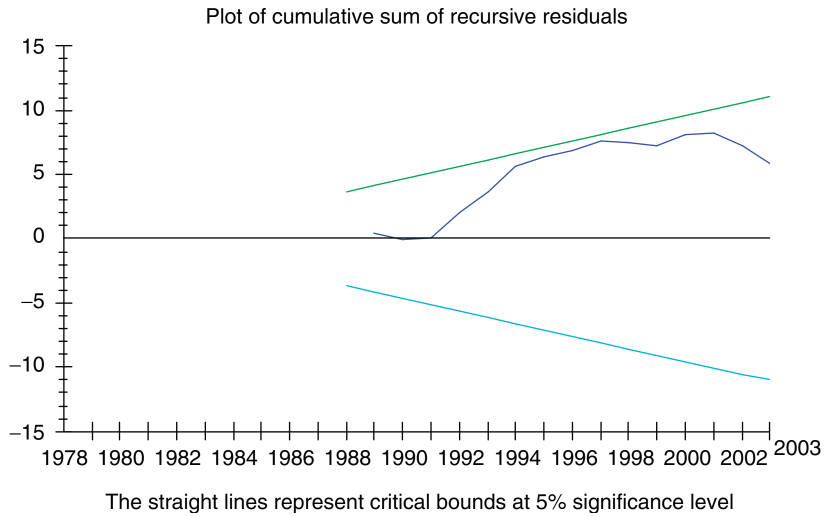

Fig. 1. ***CUSUM* test for equation 12**

The interest rate coefficients are also insignificant. We shall disregard the estimates of GH-II, GH-III and GH-IV because as Smith (2000) and Rao (2006) have pointed out, statistical techniques are only tools to summarize facts and may not answer questions of economic theory.

Therefore, we shall use the residuals from GH-I to estimate the short run dynamic equation for the demand for money with the error-correction adjustment model (ECM).

The short run ECM model is developed by using the LSE- Hendry General to Specific (GETS) framework in the second stage. Here $\Delta\ln M_t$ is regressed on its lagged values, the current and lagged values of $\Delta\ln Y_t$ and Δr_t and the one period lagged residuals from the cointegrating vector from GH-I. We have used lags up to 4 periods and using the variable deletion tests in Microfit 4.1 arrived at the following parsimonious equation:

$$\begin{aligned}\Delta\ln M_t = {} & \underset{(0.89)}{0.101} - \underset{(3.93)*}{1.337}\,ECM_{t-1} - \underset{(1.66)}{2.380}\,\Delta\ln Y_t \\ & + \underset{(3.12)*}{5.116}\,\Delta\ln Y_{t-1} + \underset{(3.35)*}{4.143}\,\Delta\ln Y_{t-2} \\ & - \underset{(2.72)*}{3.921}\,\Delta\ln Y_{t-3} - \underset{(4.64)*}{6.202}\,\Delta\ln Y_{t-4} \\ & + \underset{(3.53)*}{0.065}\,\Delta r_{t-3} + \underset{(3.40)*}{0.762}\,\Delta\ln M_{t-1} \\ & + \underset{(3.28)*}{0.702}\,\Delta\ln M_{t-2} + \underset{(1.85)**}{0.224}\,\Delta\ln M_{t-3} \qquad (12)\end{aligned}$$

$$\bar{R}^2 = 0.455, \quad SER = 0.075, \quad Peroid:\ 1978\text{-}2003$$

$$?^2_{sc} = 0.609(0.44), ?^2_{ff} = 1.408(0.24), ?^2_{n} = 0.731(0.69), ?^2_{hs} = 1.549(0.21)$$

where the absolute t-ratios are in the parentheses below the coefficients and * and ** indicates significance at the 5% and 10% level, respectively. All the estimated coefficients are significant at conventional levels except, $\Delta\ln Y_t$ is significant at about 11%. The lagged error correction term (ECM_{t-1}) has the expected negative sign implying negative feedback mechanism. That its coefficient is more than unity does not matter because it has the expected negative sign and may cause cyclical, instead of smooth adjustment towards equilibrium. The summary $?^2$ test-statistics, with p-values in the parentheses, indicate that there is no serial correlation ($?_{sc}{}^2$), functional form misspecification ($?_{ff}{}^2$), nonnormality ($?_{n}{}^2$) and heteroscedasticity ($?_{hs}{}^2$) in the residuals. Therefore, Equation 12 is well-determined.

We proceed further to test for the stability of the money demand function. When we subjected the Equation 12 to *CUSUM* and *CUSUMSQ* stability tests, neither the *CUSUM* nor the *CUSUM SQUARES* showed any instability. This implies that demand for narrow money is temporally stable in Bangladesh and therefore following Poole (1970), it can be said that money supply is the appropriate monetary policy instrument for the Central bank of Bangladesh. The plots of the *CUSUM* tests are given in Figs 1 and 2 below.

V. Conclusion

In this article, we have used time series approach and the GH technique for structural breaks to estimate the demand for real narrow money for Bangladesh for the period 1973 to 2003. Our study reveals that there exists a cointegrating relationship between real

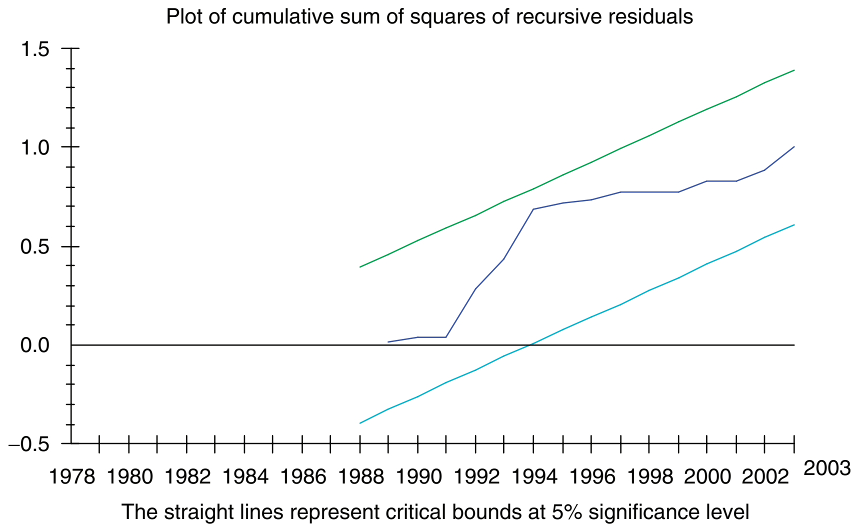

Fig. 2. ***CUSUM SQUARES*** **test for equation 12**

narrow money, real income and nominal rate of interest after allowing for structural breaks. However, of the four possible structural breaks, the one with an intercept shift in 1989 yields meaningful cointegrating coefficients. Our estimates imply that there is a well-determined and stable demand for money in Bangladesh from 1988 to 2003 and perhaps following the financial reforms in the 1980s, demand for narrow money has declined by a small amount. This result is to be expected because financial reforms improve the efficiency with which money is used in transactions.

The estimated income and interest rate elasticities are well determined and their signs and magnitudes are consistent with *prior* expectations. Our results show that income elasticity is around unity and the interest rate elasticity is negative and significant. Thus, there is no evidence that the money demand function for Bangladesh has become unstable due to deregulation and financial reforms of 1980s. Therefore, we may conclude that money supply is the appropriate monetary policy instrument to be targeted by the Central Bank of Bangladesh.

Some limitations of our study are as follows. Our specification is simple and it is desirable to add additional explanatory variables like the expected rate of inflation. However, we found that the rate of inflation is a $I(0)$ variable and therefore it is necessary to use the bounds test approach popularized by Bahmani-Oskooee and Rehman (2002). But, there is no cointegration test for this technique with structural breaks.[5] Next, as a referee has suggested it is also desirable to experiment with alternative definitions of the variables. We hope that our work would be useful for further extended work on the demand for money of Bangladesh and other countries.

Acknowledgements

We thank a referee and Rup Singh of University of the south Pacific for helpful comments.

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[5] Readers of this journal may have noted that there have been some unsubstantiated claims on the existence of small sample critical values for the bounds test. Therefore, we wish to bring to the attention of those using the bounds test that Turner (2006) has recently computed sample size adjusted critical values for the bounds test.

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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).

Notes:

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

(2) Data are available for replication on request.