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## Article

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## GENERAL & APPLIED ECONOMICS | RESEARCH ARTICLE

# On the stability of the money multiplier in Nigeria: Co-integration analyses with regime shifts in banking system liquidity

K. Moses Tule<sup>1</sup> and O. Taiwo Ajilore<sup>1\*</sup>

**Abstract:** The study hypothesized the existence of regime shifts in the conduct of monetary policy, occasioned by changing liquidity conditions in the domestic banking system in Nigeria. Within the context of this prognosis, the study tests the stability of the money multiplier, utilizing methodological procedures that allow for the explicit consideration of regime shift bias in the specification of the model and the empirical estimation. The study found the existence of a stable long run relationship between broad money and the monetary base, confirming that the necessary condition for monetary control within a multiplier framework is satisfied for Nigeria. Also, the spate of quantitative easing by the Central Bank of Nigeria to ameliorate adverse liquidity conditions and the lingering effects of the global financial crises occasioned a structural break in monetary policy, determined endogenously to have occurred in November 2009.

**Subjects:** Banking; Econometrics; Economic Forecasting

**Keywords:** money multiplier; banks liquidity; regime shifts

**Jel Codes:** E52; G21; C62



K. Moses Tule

### ABOUT THE AUTHORS

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### PUBLIC INTEREST STATEMENT

A prerequisite for policy regimes that target the levels of monetary stocks is the ability of the monetary authority to control the quantity of money stock supplied and forecast the changes in the factors that affect the resulting money supply. The foregoing requires a long run stable relation between base money and the measure of money stock. The need then exists to empirically verify the existence of this condition to provide credence for the money targeting framework of monetary policy in Nigeria. Existing empirical research toward this end for Nigeria has been found to be weak in light of recent innovations in econometric analysis of unit root and co-integration tests. This study obviates these weaknesses and empirically validates a stable long run money multiplier relation for Nigeria. This provides the needed credibility for the monetary targeting policy regime of Central Bank of Nigeria.

## 1. Background to the study

A monetary policy regime that targets the money stock owes its effectiveness to the stability of the money multiplier relation; otherwise, the primary objective of monetary policy, i.e. price stability, cannot be achieved by targeting the monetary stock (Feige & Parkin, 1971). As it were, the subject of stability of the money multiplier has attracted phenomenal attention in empirical literature.

Authors such as Bomhoff (1977), Johannes and Rasche (1981), Chitre (1986), Nachane and Ray (1989), and Ray and Madhusoodan (1992) have employed the time series forecasting techniques to assess the empirical predictability of the money multiplier. The increasing advancement and popularity of co-integration analysis have further provoked examining the question in terms of co-integrating relation between given monetary aggregate and high-powered money. Nachane (1992), Hansen and Kim (1995), Ford and Morris (1996), Baghestani and Mott (1997), Sen and Vaidya (1997) and Darbha (2002) present examples of authors that test for the stability of a long run relation between money stock and reserve money (RM) within the framework of co-integration analyses.

The weak point of most of the co-integration analyses of money multiplier stability research is their application of standard ADF and PP classes of co-integration tests. Gregory and Hansen (1996) have argued that since these tests presume that the co-integrating vector is time invariant under the alternative hypothesis, they have low power in detecting a discrete change in a policy regime which may lead to a one-time shift in the parameter vector.

These issues become relevant for Nigeria, as monetary targeting remained the framework for the implementation of monetary policy by the Central Bank of Nigeria since 1974. Given that the question whether the money multiplier is stable and predictable provides the theoretical plank for the adoption of this monetary strategy, a good measure of research inquiries have sought to provide definitive answer. (see Doguwa (1994 and), Doguwa (1997) and Uchendu (1995)).

These studies have mainly relied on the use of univariate tests applied on the money multiplier in order to assess its stability, as well as testing the hypothesis of predictability of the money multiplier based on bivariate models linking the monetary base to money supply. These empirical procedures have nonetheless relied heavily on the assumption of the existence of a prior and potential common trend between base money and money supply.

This assumption is however violated in the presence of marked regime shifts in liquidity conditions of the domestic money market, particularly in the banking sector, since the use and effectiveness of the required reserves ratio, an important tool to regulate the money creation capacity of commercial banks, is predicated on the liquidity conditions of the banking system. The banking system in Nigeria has transverse switching regimes of liquidity surfeit and liquidity crunch that have occasion switches between tight and easy monetary policy stance by the Bank.

Monetary management prior to the global financial crisis focused on mopping up excess liquidity from the banking system mainly using open market operations (OMO). However, following the global financial crisis and the post-crisis period, monetary policy has focused on managing liquidity shortfalls, thus demanding liquidity injections into the system. However, since November 2010, monetary policy stance has shifted back to monetary tightening because of inflationary pressures arising from the anticipated and subsequent injections by the Asset Management Corporation of Nigeria (AMCON) and the residual effects of previous CBN's monetary easing measures.

These changes in banking system liquidity conditions highlight the transition from one regime of managing liquidity conditions to another. It is then necessary that such transition be accounted for in the long run tests of stability of the money multiplier. The negligence of the existence of this phenomenon undermines the plausibility of existing empirical evidence that links the monetary base to money supply without paying attention to the fact of regime shift in banking sector liquidity conditions over time. The key empirical innovation of this study over previous empirical investigations on



the stability of Nigeria's money multiplier lies in the introduction of the reality of regime shift in banking sector liquidity conditions in the specification of the model and the empirical estimation.

The rest of the paper proceeds as follows: Following this introduction, Section 2 reviews previous literature on the stability of the money multiplier in Nigeria. Section 3 articulates the framework of theory of the money multiplier model of money supply, while Section 4 presents the empirical model and the econometric procedures for its estimation. Section 5 discusses the results while Section 6 concludes.

## 2. Review of extant literature

The Central Bank of Nigeria Act 1958 and subsequent reenactment and amendments thereto assigned the promotion of monetary and price stability and a sound financial system as core mandates of the Bank. To this end, the Bank is expected to design appropriate monetary policy that would lead to the achievement of the set objectives. In pursuance of this mandate, monetary policy by the Bank over the years has essentially focused on influencing the growth in money supply that is consistent with the core objectives of price and exchange rate stability, real growth, and overall soundness and stability of the financial system. The Bank's ultimate targets are pursued through the adoption of money supply, in particular broad money (M2) as the intermediate target variable and base money (RM) as the operational target.

The reserve or base money which is a liability to the Bank consists of currency in circulation plus DMBs' reserves with the central bank. The strategy of monetary aggregate targeting involves the regulation of the amount of base money through policy actions designed to influence its growth path, largely through OMO. The transmission mechanism thus proceeds from the base money stock, and moves toward inflation, passing through the money supply aggregate, M2. The overall objective of this framework is to keep money supply close to its estimated target level, preventing significant excess or shortfall that may lead to precarious deviations in outcomes of inflation and economic growth compared with their targets.

The monetary targeting establishes the linkages between central banks' monetary policy actions to affect base money and influences the broad money stock, M2. Inherent in this model is the role of the money multiplier, defined as the ratio of the money supply to base money. Two assumptions about the money multiplier are crucial for the effectiveness of the monetary targeting framework as a strategy for monetary policy. These are the stability and predictability of the money multiplier. The former involves the stationary properties of the money multiplier, while the latter entails the existence of stable long run relationship between the monetary base and money supply.

An important assumption underlying the efficacy of the monetary aggregates as a policy target variable is the existence of a stable long run relationship between the monetary aggregates such as M2 and RM. Consequently, the issue of examining whether the money multiplier is stable and predictable for Nigeria has attracted a good measure of research inquiry for the monetary targeting policy framework in Nigeria.

Doguwa (1994) undertook the earliest empirical effort at examining the issue of stability of the money multiplier and its implications for the conduct of monetary policy for Nigeria. The study utilized a number of eclectic parametric and non-parametric approaches to conclude that the Central Bank of Nigeria's policy actions were independent of the adjusted money multiplier.

Uchendu (1995) investigated the nature of the relationship between money supply and base money in order to establish the veracity of the indirect approach to monetary control in Nigeria, utilizing annual data for the period 1976–1993. The study relied on a moving average regression technique on a combined regression and time series model of money stock and base money to test for structural breaks. The study found evidence for the existence of an unstable relationship between the money stock and base money, in line with Doguwa (1994), violating the theoretical basis for the use

of indirect approach to liquidity management. By this result, Nigeria had no theoretical plank for migrating to the indirect monetary targeting.

Doguwa (1997) revisited the question of the stability of the money multiplier, following Uchendu (1995) to show how the adjusted base money acts as the main link to money supply. The study relied on the method provided by the forward recursive cusum of squares that uses a confidence interval which is distributed as Pyke's-modified Kolmogrov-Smirnov statistic for the monthly data spanning from January 1991 to December 1995. The study concluded that the functions underlying the adjusted money multipliers were not stable for most of the period of the study, further confirming his earlier findings.

We argue that these studies have mainly relied on the use of univariate tests applied on the money multiplier in order to assess its stability and that the apparent rejection of the existence of stable long run relationship in the money multiplier may not be unrelated to the neglect of the fact of significant discrete changes in the Nigeria's financial system and the conduct of monetary policy in their empirical procedures. Hence, this study considers that the model with endogenous structural breaks may more aptly underpin efficiently the relations between broad and base money in Nigeria.

### 3. Money multiplier: framework of theory

The reserve or base money which is a liability to the Bank consists of currency in circulation plus DMBs' reserves with the central bank. The strategy of monetary aggregate targeting involves the regulation of the amount of base money through policy actions designed to influence its growth path, largely through OMO. The transmission mechanism thus proceeds from the base money stock, and moves toward inflation, passing through the money supply aggregate, M2. The overall objective of this framework is to keep money supply close to its estimated target level, preventing significant excess or shortfall that may lead to precarious deviations in outcomes of inflation and economic growth compared with their targets.

Monetary targeting establishes the linkages between central banks' monetary policy actions to affect base money and influence the broad money stock, M2. Inherent in this model is the role of the money multiplier, defined as the ratio of the money supply to base money. Two assumptions about the money multiplier are crucial for the effectiveness of the monetary targeting framework as a strategy for monetary policy. These are the stability and predictability of the money multiplier. The former involves the stationary properties of the money multiplier, while the latter entails the existence of stable long run relationship between the monetary base and money supply.

The notion of fractional reserve banking explains the determinants of the money stock in terms of the money multiplier and the monetary base or the so-called high-powered money. The formal relation between the money supply and high-powered money is indicated as:

$$M = D + C \quad (1)$$

where the money supply,  $M$ , consists of deposits of commercial banks,  $D$ , and currency held by the public,  $C$ . The monetary base,  $B$ , on the other hand is made up of currency held by the public,  $C$ , and the sum of the required reserves (RR) and excess reserves (ER) of the domestic banking system:

$$B = C + RR + ER \quad (2)$$

The relation between money stock,  $M$ , and the base money,  $B$ , is expressed as the ratio of  $M$  and  $B$ :

$$\frac{M}{B} = \frac{D + C}{C + RR + ER} \quad (3)$$

By dividing the right-hand side of Equation (3) by  $D$ , we obtain the relations:

$$\frac{M}{B} = \frac{1 + \frac{C}{D}}{\frac{C}{D} + \frac{RR}{D} + \frac{ER}{D}} \quad (4)$$

But  $\frac{C}{D}$  is the cash deposit ratio,  $c$ ,  $\frac{RR}{D}$  is the required reserves ratio  $r_r$  and  $\frac{ER}{D}$  is the excess reserves ratio  $r_e$ . Substituting into Equation (4):

$$\frac{M}{B} = \frac{1 + c}{c + r_r + r_e} \quad (5)$$

Thus, the money stock,  $M$ , becomes:

$$M = \frac{1 + c}{c + r_r + r_e} B \quad (6)$$

Equation (6) expresses money stock in terms of the base money and the money multiplier. The quotient of Equation (6) is the money multiplier,  $m$ .

$$m = \frac{1 + c}{c + r_r + r_e} \quad (7)$$

The equation indicates that the higher the supply of base money, the higher the money supply. The size of the money multiplier is determined by the behavior of the banking public in their preference for holding cash relative to bank deposits, ( $c$ ); the monetary policy stance of the central bank in determining the required reserves ratio ( $r_r$ ); and the incentives for holding ER by the commercial banking system, ( $r_e$ ). The lower these ratios are, the larger the money multiplier is.

The introduction of the money multiplier has shown the link between the balance sheet position and monetary aggregate as a first step, in relation to prices and the real economy. The required reserves ratio is an important tool on the part of central banks to regulate money creation by commercial banks. While some economists do not consider ER in determining base money, the monetarists give more importance to the same. According to this view, banks, like other businesses, are subject to uncertainties inherent in banking operations, as such they keep ER to hedge these risks. The size of ER they hold is subject to the costs of such holding, the most significant being the market rate of interest forgone in holding ER. Since the money supply is inversely related to excess reserves ratio, decline in the excess reserves ratio of banks tends to increase money supply.

If the money multiplier,  $m$ , is fairly stable, the central bank can manipulate money supply ( $M$ ) by manipulating the monetary base, utilizing the instruments of monetary policy at its disposal. However, the stability of the money multiplier is in turn dependent on its determinants, namely: the currency ratio and the required and excess reserves ratios. Since these ratios are liable to change, the money multiplier is volatile in the short run. Thus, one of the major issues in monetary policy is the stability and predictability of the money multiplier; empirical verifications of such long-term stability have produced conflicting outcomes.

#### 4. Empirical model and estimation procedures

Combining the money stock,  $M$ , and money multiplier,  $m$ , of Equations (6) and (7), respectively, the money multiplier relations is written in the deterministic form as:

$$M_t = mB_t \quad (8)$$

where all notations remained as defined. The logarithmic transformation of Equation (8) yields the general form as:

$$\ln M_t = \mu + \alpha \ln B_t + \varepsilon_t \quad (9)$$

where  $\ln$  indicates the natural logarithmic transformation of the respective variables and  $\varepsilon_t$  is the usual white noise error term. The intercept of Equation (9),  $\mu$ , corresponds to the money multiplier if the proportionality relation  $\alpha = 1$  holds. Thus, the stability and predictability of the money multiplier model require that,  $\ln M$  and  $\ln B$  are co-integrated, with coefficient restrictions such that  $\mu = 0$  and  $\alpha = 1$ .

A vast majority of empirical verification of the stability of the money multiplier have relied on the use of conventional residual-based Engle–Granger’s ADF and PP classes of tests in the sense of a linear combination of non-stationary variables being stationary. However, Gregory and Hansen (1996) emphasized that while two non-stationary series may be co-integrated in the sense of their linear combination, there is the possibility that the co-integrating vector (linear combination) has shifted at one unknown point in the sample. In this instance, the cointegrating vector is no longer time invariant, and the standard tests for co-integration are not appropriate.

In light of the foregoing, and given a clear evidence of regime switch in banking sector liquidity conditions within the sample period, a model of money supply function that ignores such regime switches, in the view of Ford and Morris (1996), is indicative of misspecification in the money supply function. Existing empirical verifications of the stability of money multiplier function for Nigeria suffer from this empirical misspecification; hence, the current study employs an empirical procedure that obviates this weakness.

Consequently, our empirical procedure draws on Gregory and Hansen (1996), by testing for the null of no co-integration between our measure of money stock (M2) and RM, against the alternative of co-integration with a regime shift at an unknown point in the sample.

We estimate Equation (9) employing the method of Gregory and Hansen (1996). The method modifies the conventional Engle and Granger (1986) approach for testing the existence of a single co-integrating relationship by allowing the coefficients to be subject to a structural break. The timing of the break is also data based, rather than imposing the break date on the data. We adopt the most general regime shift model that permits shifts in both the intercept and slope coefficients, corresponding to the level and regime change models of Gregory and Hansen (1996). Specifying Equation (9) in this form yields:

$$\ln M_t = \mu_1 + \mu_2 \varphi_{t\tau} + \alpha_1 \ln B_t + \alpha_2 \ln B_t \varphi_{t\tau} + \varepsilon_t, \quad t = 1, \dots, n \quad (10)$$

$\mu_1$  and  $\mu_2$  are, respectively, the intercept before the shift and the change in the intercept at the time of the shift.  $\alpha_1$  and  $\alpha_2$  denote, respectively, the co-integrating slope coefficients before the regime shift and change in the slope coefficient at the time of the shift.

$\varphi_{t\tau}$  is a dummy variable to indicate the break date defined as:

$$\varphi_{t\tau} = \begin{cases} 0 & \text{if } t \leq [n\tau], \\ 1 & \text{if } t > [n\tau] \end{cases}$$

$\tau$  is an unknown parameter within the domain (0, 1) used to model the location of the structural break. The technique comprises sequential search of the break date within the permitted range of  $[0.15n] \leq t \leq [0.85n]$ . For each possible break date, a statistic to test the null of no co-integration against the alternate co-integration with a structural break is computed.

The critical values and the associated asymptotic distributions of the test statistics are contained in Gregory and Hansen (1996) as follows:



$$\begin{aligned} Z_{\alpha}^{*} &= \inf_{\tau \in T} Z_{\alpha}(\tau) \\ Z_t^{*} &= \inf_{\tau \in T} Z_{\alpha}(\tau) \\ ADF^{*} &= \inf_{\tau \in T} ADF_{\alpha}(\tau) \end{aligned}$$

If the minimum value obtained for the test statistic exceeds the Gregory and Hansen critical value, the null of no co-integration is rejected for the alternate of co-integration with structural break. It should be noted that the Gregory-Hansen approach to testing co-integration assumes null of co-integration. There are tests that assume the null of no co-integration in the spirit of Pesaran and Shin (1996). To this end, we implement the Arai-Kurozumi co-integration test (Arai & Kurozumi, 2007) that uses the LM-type test statistic assuming the null of no co-integration. It thus provides a robustness check to the Gregory-Hansen approach.

Before testing the presence of a long run relationship between base money and M2 balances in the presence of regime shifts, we have to determine the stationary properties of variables in the model. Due to our prognosis of possible existence of structural changes that might have occurred in the period under study, we prefer the procedure developed by Zivot and Andrews to test the null of unit root against the break stationary alternative hypothesis to the conventional Augmented Dickey and Fuller (1979) test which has remained the workhorse of test for stationarity. Zivot and Andrews (1992) endogenous structural break test is a sequential test which utilizes the full sample and uses a different dummy variable for each possible break date. The break date is selected where the  $t$ -statistic from the ADF test of unit root is at a minimum (most negative). Consequently, a break date will be chosen where the evidence is least favorable for the unit root null. The critical values in Zivot and Andrews (1992) are different to the critical values in Perron (1989). The difference is due to the fact that the selection of the time of the break is treated as the outcome of an estimation procedure, rather than predetermined exogenously.

To keep in consonance with the form of Gregory and Hansen co-integration analysis as indicated in Equation (10), we estimate the Zivot and Andrews' model to test for a unit root against the alternative of a one-time structural break in the data. Consequently, we specified the test model of the form:

$$\Delta y_t = c + \alpha y_{t-1} + \beta t + \theta DU_t + \gamma DT_t + \sum_{j=1}^k d_j \Delta y_{t-j} + \varepsilon_t \quad (11)$$

where  $DU_t$  and  $DT_t$  are indicator dummy variables, respectively, for a mean shift and the corresponding trend shift variable occurring at each possible break date ( $\tau$ ). Formally,

$$DU_t = \begin{cases} 1 & \text{if } t > \tau \\ 0 & \text{otherwise} \end{cases} \text{ and } DT_t = \begin{cases} t - \tau & \text{if } t > \tau \\ 0 & \text{otherwise} \end{cases}$$

The model tests the null hypothesis  $\alpha = 0$ , that the series  $y_t$  contains a unit root with a drift that excludes any structural break, against the alternative hypothesis  $\alpha < 0$ , that the series is a trend stationary process with a one-time break occurring at an unknown point in time (see details in Zivot and Andrews (1992)). We also employ the alternative strategy to testing unit root under structural break as suggested by Harvey, Leybourne, and Taylor (2013). This strategy brings to bear the prior information on the estimation of the break dates. Since it is widely acknowledged that the global financial system was adversely affected in 2008, leading to one of the major recessions in recent years, we will use this information while estimating the break in the relationship.

## 5. Results and discussion

The study employed monthly observations spanning the periods 2003(01)–2015(08) for Nigeria, yielding 152 observations. The choice of period was informed by consistent availability of high frequency data that are considered suitable for the type of analysis intended in the study. The period

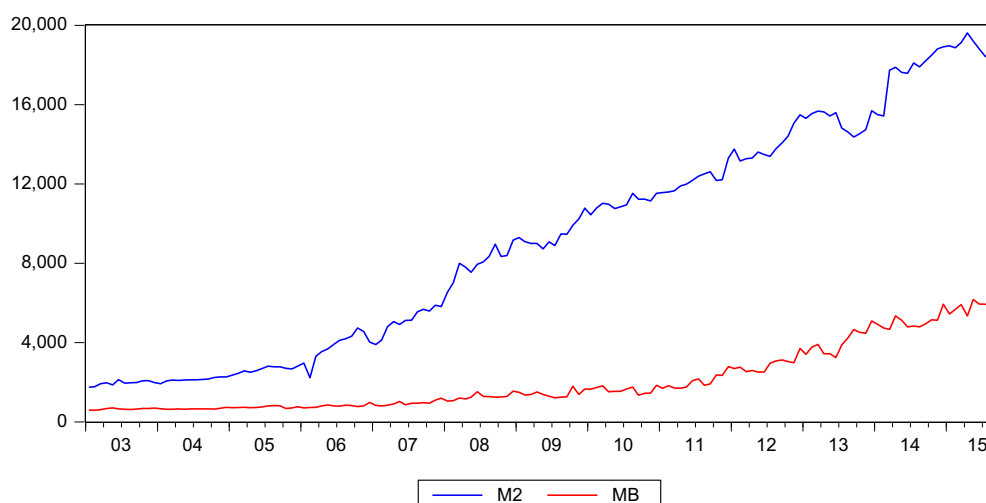
also covers the two regimes of surfeit and squeeze in liquidity positions in the domestic banking system as we encapsulate in the empirical procedures.

The variables employed for the study consist of base money and broad money (M2). The data were sourced from the online Database of Statistics, Central Bank of Nigeria. A preliminary graphical examination of the two series is shown in Figure 1. An irregular pattern of changes is clearly observable in the time path of the plots of the two series.

According to Driffill and Sola (1998), a way to overcome this in detecting more reliably the stationarity patterns in series with such behavior is to allow the series to switch between two regimes.

Since Perron (1989), it is well known that in the presence of structural break, the standard ADF statistic is no longer powerful and is biased toward a false unit root null when the data are trend stationary with a structural break. Thus, unit root based on the ADF statistic might not necessarily suggest non-stationarity but rather that there is a break. To detect this, we will juxtapose the results based on the standard ADF and the Zivot and Andrews' unit root tests. We proceeded along this line by implementing the Zivot and Andrews' unit root test of the version indicated in Equation (11) with the inclusion of  $\beta_t$  term. This version becomes most appropriate, given the fact that all the variables indicate an upward trend as Figure 1 indicates. Ben-David and Papell (1997) submit that if a series exhibits a trend, then estimating the model without trend may fail to capture some important characteristics of the data.

**Figure 1. Plots of base money (mb) and broad money (m2).**



**Table 1. The Perron and Yabu (2009) test for structural break**

Series	Model	$W_{RQF}$	Break
LMB	Intercept	3.313	2012M06
	Slope	1.011	2010M09
	Both	7.909	2011M09
LM2	Intercept	23.152	2007M02
	Slope	2.030	2009M10
	Both	38.030	2008M01

However, before testing the presence of unit roots in the series, it is a good idea to pretest the possibility of break in the data. To this end, we conduct a pretest for discontinuity in each of the series, employing the Perron–Yabu test (Perron & Yabu, 2009), as indicated in Table 1 above.

We test the null hypothesis of no structural change in the intercept, slope, and both using the Perron and Yabu’s (2009) robust feasible quasi-generalized least squares (FGLS) Wald statistic, WRQF. The WRQF is a Wald statistic based on least squares estimation of Cochrane–Orcutt transformation of a scalar random variable, using a modified estimate of  $\alpha$ . The results in Table 1 indicate that there is discontinuity in the data. Given these prognostic results of possible structural changes, we are now in the position to determine the stationary properties of variables in the model, while accounting for structural breaks in the data-generating process.

Panel A of Table 2 presents the results for Zivot–Andrews unit root tests on log-levels and log-differences of broad money (M2) and base money (MB). These results clearly indicate we can reject the null of unit root for both series at a 1% significance level, thus confirming they are both I(1) series. Figure 2 further identifies endogenously the point of the single most significant structural break (TB) in the time series. The break date for each time series is reported in Table 2 as 2006M03 and 2013M07, respectively, for M2 and MB. Given the prior knowledge of the possible break dates, especially during the 2008 global recession, and to account for the delayed effects in the structural changes during the period, we further consider the method suggested in Harvey et al. (2014). This serves as a robustness check on the existence of break in the data-generating process for each of the variables. One other advantage of this approach is that it obviates the problem associated with small sample bias, given the relatively short sample size. Results of unit root test using the method advanced by Harvey et al. (2014) are indicated in Panel B of Table 2. This further confirmed both series to be I(1) processes.

The confirmation of the existence of structural breaks in the integrating properties of the time series employed in the study justifies the choice of the G–H co-integrating tests employed in the study. According to Valadkhani, Layton, and Pahlavani (2005), not taking into consideration the presence of structural breaks may generate biased empirical results.

**Table 2. Unit root with a structural break in both the intercept and trend**

<b>Panel A: Zivot–Andrews</b>		
<b>Variables</b>	<b>t-statistics*</b>	<b>Break point</b>
<i>Log-levels</i>		
M2	–2.37211	2008M03
MB	–3.22113	2013M07
<i>Log-differences</i>		
M2	–10.52881	2006M03
MB	–9.945609	2013M07
<b>Panel B: Harvey–Leybourne–Taylor (2013)</b>		
<b>Variables</b>	<b>MDF statistics</b>	<b>Break point</b>
<i>Log-levels</i>		
M2	–2.371567	2009M08
MB	–3.280699	2008M12
<i>Log-differences</i>		
M2	–8.178207	2008M12
MB	–6.915647	2009M01

\*The critical values for Zivot–Andrews test are –5.57, –5.08, and 4.82 at 1%, 5%, and 10% levels of significance, respectively, while the critical values for Harvey–Leybourne–Taylor are –4.02, –3.49, and –3.20 at 1%, 5%, and 10% levels of significance, respectively.

**Figure 2. Plots of estimated timing of structural breaks.**

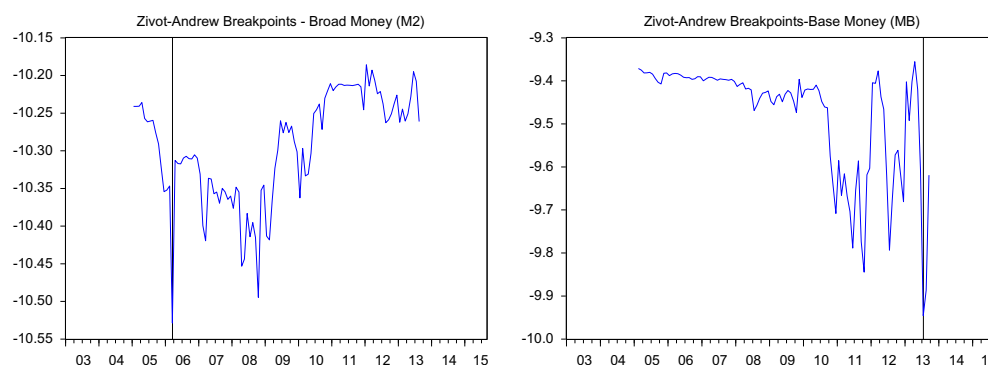


Table 3 presents the results of the G-H tests of co-integration. The three test statistics are statistically significant at the 5% level, thereby supporting the existence of co-integration with a structural break or regime shift. In addition, the break date is detected to be November 2009. This period aptly corresponds with the time monetary policy stance had to switch from being tightened to being accommodative in order to ameliorate the debilitating effects of the global financial crisis on the domestic financial system, in particular the banking sector. According to Piehl et al. (1999), knowledge of break point has important implications, as it is central for accurate evaluation of any program intended to bring about structural changes.

As a further test of robustness of the G-H test of co-integration, Table 4 presents the results from the co-integration analysis with the approach suggested in Arai and Kurozumi (2007). The Arai-Kurozumi approach compliments the G-H test. While the G-H is a residual-based test for the null of no co-integration against the alternative of co-integration with a structural break of unknown timing, the Arai-Kurozumi approach is a test for the null hypothesis of co-integration with a structural break against the alternative of no co-integration, utilizing the Lagrange multiplier (LM)-based test statistic.

**Table 3. Gregory-Hansen co-integration test results**

Model specification	Testing procedure	Test statistic	5% critical value*	Break point
<i>Level and regime shifts</i>				
	ADF*	-7.95**	-6.00	2009M11
	$Z_t^*$	-7.96**	-6.00	2009M11
	$Z_a^*$	-89.43**	-68.94	2009M11

\*Critical values were obtained from Gregory and Hansen (1996).

\*\*Statistical significance at the 5% level.

**Table 4. Arai-Kurozumi co-integration test results for level and regime shifts**

Break date	2009M04
Break fraction	0.421
L4	0.134*
L12	0.069
La	2.985**
P(L4)	0.058
P(L12)	0.045
P(La)	2.613**

Note: P(z) refers to prewhitened statistics.

\*Level of significance at 5%.

\*\*Level of significance at 1%.

At the same time, the break date suggested by the Arai-Kurozumi approach (Table 4) is also justifiable since the effects of the global recession are widely acknowledged to begin to bottom out in 2009.

Next, we test for exogeneity and co-breaking between M2 and MB. In view of the global financial crisis that rocked the financial world in 2008, the possibility of simultaneous decline of both M2 and MB cannot be ruled out, which could lead to co-breaking in the relationship. The basic idea of co-breaking states that series subjected to breaks individually could nevertheless exhibit stability if they are linearly combined, in much the same way that co-integration guarantees the stationarity of non-stationary series when linearly combined. To investigate this phenomenon, we test for super exogeneity and the possibility of co-breaking in the model following the work of Ericsson, Hendry, Mizon, and Grayham (1998), Hendry and Ericsson (1991), and Beyer (1998). Given the bivariate model under investigation, we first estimate the co-integrated VAR model to factorize the model into conditional and marginal models. As the model is bivariate, we only expect at most one potential co-integrating vector. The conditional-marginal factorization procedure can be stated as:

$$\prod_{t=1}^T D(LM2_t, LMB_t; \theta) = \prod_{t=1}^T D(LM2_t | LMB_t, X_{t-1}, \phi_1) D(LMB_t | X_{t-1}, \phi_2)$$

where  $D(LM2_t, LMB_t; \theta)$  is the joint distribution for the two variables,  $D(LM2_t | LMB_t, X_{t-1}, \phi_1)$  is the distribution for the conditional model, and  $D(LMB_t | X_{t-1}, \phi_2)$  is the distribution for the marginal model, with  $\phi = f(\theta)$ . For weak exogeneity, the parameters in the data-generating processes for these variables have to be variation free. Super exogeneity implies that the variations in  $\phi_2$  do not induce variations in  $\phi_1$ . To establish weak exogeneity of LMB, we follow the literature estimating a co-integrated VAR of the form:

$$\begin{pmatrix} \Delta LM2_t \\ \Delta LMB_t \end{pmatrix} = \begin{pmatrix} \delta_1 \\ \delta_2 \end{pmatrix} + \begin{pmatrix} \alpha_1 \\ \alpha_2 \end{pmatrix} (\beta_1, \beta_2) \begin{pmatrix} LM2_{t-1} \\ LMB_{t-1} \end{pmatrix} + \begin{pmatrix} \Gamma_{11} & \Gamma_{21} \\ \Gamma_{12} & \Gamma_{22} \end{pmatrix} \begin{pmatrix} \Delta LM2_{t-1} \\ \Delta LMB_{t-1} \end{pmatrix} + \begin{pmatrix} \varepsilon_{1t} \\ \varepsilon_{2t} \end{pmatrix}$$

If  $\alpha_2 = 0$ , then LMB is weakly exogenous, and if, in addition,  $\Gamma_{12} = 0$ , then LMB is strongly exogenous. A test for super exogeneity is not rejected if none of the shifts in parameters of the marginal model affects the stability of the conditional model (Beyer, 1998). To test for weak exogeneity between these variables, we employ Johansen's (1998) co-integration analysis. The co-integration test suggests that both trace and max-eigen statistics are not significant at conventional levels. However, based on economic reasoning, we retain the first relation. The test for strong exogeneity shows that  $\chi^2(4) = 9.73$  with  $p$ -value of 0.045, suggesting that MB is strongly exogenous. It remains to prove super exogeneity of MB, which involves establishing the stability of the conditional model.

Following Beyer (1998) and given the results above, the conditional model of interest then reduces to the augmented distributed lag (ADL) model of the form:

$$\Delta LM2_t = \sum_{i=1}^{12} \omega_i \Delta LM2_{t-i} + \sum_{i=1}^{12} \eta_i \Delta LMB_{t-i} + \alpha + \phi LM2_{t-1} + \kappa LMB_{t-1} + \lambda t + \varepsilon_t \quad (12)$$

which can be stepwise simplified to obtain a more parsimonious model. The results reported in Table 5 are based on a general-to-specific estimation procedure with the nominal critical value set to 5%. The marginal model is modeled as an AR process with the Durbin-Watson consistent order chosen. The maximum order of 12 is used considering the monthly frequency of the data. By adopting a Durbin-Watson consistent model, we arrive at AR(7) model with optimal DW statistic of 1.9989. The results are given in Panel B of Table 5. To prove the stability of the conditional model, we employ two approaches. We first employ the stability and structural break tests. Secondly, we do a two-step impulse saturation test, whereby we create impulse dummies for every observation in the sample. The first half of the (impulse dummies) indicators is added to our AR(7) and a general-to-specific procedure to eliminate the insignificant indicators is carried out. The next half is likewise added and the insignificant indicators eliminated. We then combine the retained significant indicators from the



**Table 5. Estimated conditional and marginal models**

**Panel A: Conditional model**

Variable	Coefficient	Prob.*
C	0.767967	0.0520
LM2(−1)	−0.046761	0.1241
LMB(−1)	−0.063456	0.0792
Trend	0.001734	0.0879
DLM2(−1)	−0.167762	0.0442
DLM2(−8)	−0.165705	0.0369
DLMB(−5)	−0.109864	0.0259

**Panel B: Marginal model**

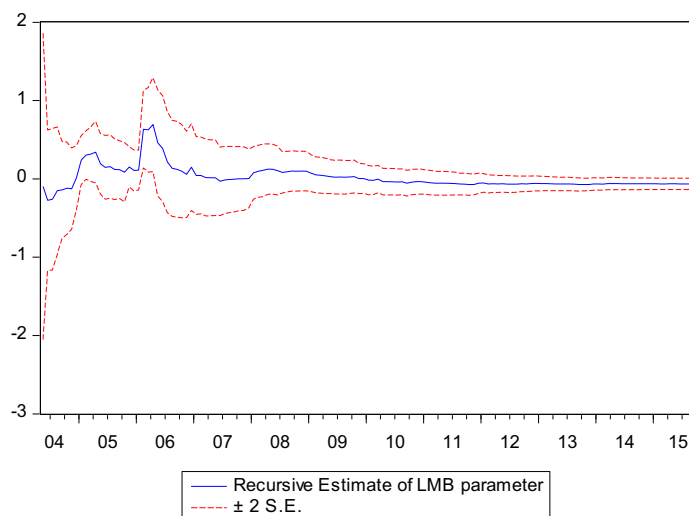
Variable*	Coefficient	Prob.
C	0.015223	0.0000
AR(1)	−0.418922	0.0000
AR(2)	−0.314910	0.0006
AR(3)	−0.294307	0.0015
AR(4)	−0.174160	0.0739
AR(5)	−0.028819	0.7293
AR(6)	−0.065572	0.4418
AR(7)	0.060159	0.5093
SIGMASQ	0.006833	0.0000

\*Optimal DW statistic is 1.9989.

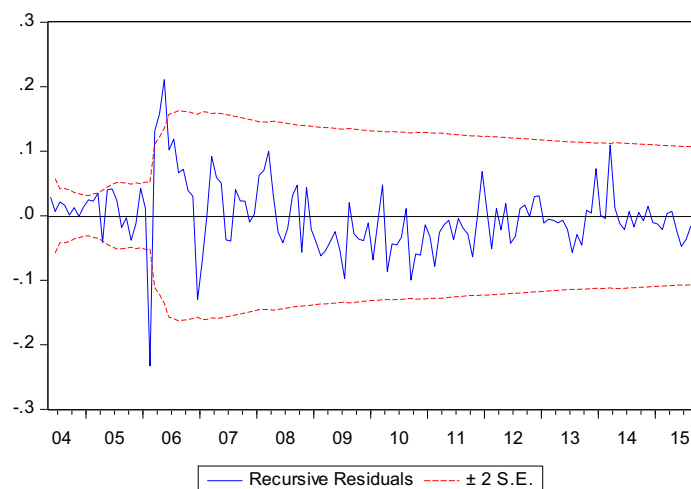
marginal models, which are then tested for significance in the conditional model given in Equation (12). Testing for the co-breaking is likewise tested where we take note of non-overlapping in timing for the instability to indicate co-breaks in both series. The plots presented in Figures 3 and 4 point to the possibility of co-breaking in 2006 and rule out the super exogeneity of LMB, even though this variable is weakly and strongly exogenous.

The Bai-Perron multiple breakpoint estimates show the possibility of structural breaks in March/April 2006 and January 2008 in the conditional model without having factored in the saturation

**Figure 3. Recursive estimate of the MB parameter in the conditional model.**



**Figure 4. Recursive residual estimate of the conditional model.**



impulse dummies. However, the stability test of the marginal model shows that the break date only occurred in April 2006. Thus, there is no clear-cut evidence supporting co-breaking, given the stability test, Table 6.

Given that the stability test does not give robust results, we proceed with the impulse saturation test to clarify the existence of co-breaking in the model. Augmenting the model with impulse dummies for sample observations and estimating the AR(7)—the marginal—model for change in log of MB, we found the significant indicators for 2004M01, 2006M01, 2006M02, 2006M03, 2006M05, 2006M06, 2006M10, 2007M03, 2007M04, 2007M07, 2007M11, 2008M01, 2008M02, 2008M03, 2008M08, 2008M09, 2008M12, and 2014M03. However, in the conditional model, only the impulse

**Table 6. Multiple breakpoint test (Bai & Perron, 2003) of the conditional model**

Sequential <i>F</i> -statistic determined breaks: 2			
<i>Break test</i>	<i>F</i> -statistic	<i>Scaled F</i> -statistic	<i>Critical value</i> **
0 vs. 1*	12.30849	86.15945	21.87
1 vs. 2*	3.636783	25.45748	24.17
2 vs. 3	1.096975	7.678826	25.13
<i>Break dates</i>	Sequential	Repartition	
1	2006M04	2006M03	
2	2008M01	2008M01	

\*Significant at the 0.05 level.

\*\*Bai-Perron (*Econometric Journal*, 2003) critical values.

**Table 7. Quandt–Andrews unknown breakpoint test for the marginal model**

<i>Statistic</i>	<i>Value</i>	<i>Prob.</i>
Maximum LR <i>F</i> -statistic (2009M06)	1.921757	0.4049
Maximum Wald <i>F</i> -statistic (2005M02)	46.15305	0.0000
Exp LR <i>F</i> -statistic	0.404148	0.8855
Exp Wald <i>F</i> -statistic	18.59905	0.0000
Ave LR <i>F</i> -statistic	0.784216	0.7352
Ave Wald <i>F</i> -statistic	17.56240	0.0083

dummies for 2006M02, 2006M03, 2007M03, 2008M01, 2008M02, 2008M03, 2008M09, 2008M12, and 2014M03 are found to be significant. Thus, co-breaking happened between M2 and MB in 2004M01, 2006M01, 2006M05, 2006M06, 2006M10, 2007M04, 2007M07, 2007M011, and 2008M08. In other words, there is no doubt that the relationship under investigation has been subjected to intense structural breaks; in most cases, however, there is evidence that the variables are dipped together, thereby cancelling out the breaks, Table 7.

The aggregates of results above summarize the existence of a long run co-integrated and stable relationship between the monetary aggregate (M2) and base money, in the presence of a regime shift, which implies a stable money multiplier relation in the monetary targeting framework of monetary policy in Nigeria. We argue that the significant departure from the findings from previous literature on Nigeria underscores the superior efficiency of our empirical procedures.

## 6. Summary and conclusion

The thrust of the paper is to examine if there exists a stable relation between a measure of the monetary aggregate (M2) and the monetary base, while admitting the existence of a regime shift in the conduct of monetary policy from an easy to tight policy stance, occasioned by changes in the liquidity conditions in the domestic banking system. This is a significant departure from existing literature on Nigeria that has attempted the same, but with empirical procedures that failed to incorporate the implications of apparent regime shifts in the conduct of monetary policy. We suggest that the inability to establish the existence of a stable money multiplier by these studies is due to the weaknesses of empirical procedures.

The verification of the existence of a stable long run relationship between broad money and the monetary base by this study has confirmed that the necessary condition for monetary control within a multiplier frame work is satisfied for Nigeria. Also, the spates of quantitative easing measures embarked upon by the Bank to ameliorate the adverse effects of the global financial system occasioned a structural break in the conduct of monetary policy which was endogenously determined to have occurred in November 2009.

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