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STOCK PRICES AND DEMAND FOR MONEY: EVIDENCE FROM NIGERIA

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ABSTRACT

This paper investigates the relationship between stock prices and the real money demand for Nigeria using unit root test and the cointegration with structural break. There are five important results for narrow and broad money in Nigeria. First, there is cointegration between narrow and broad money and their determinants, income, interest rate, exchange rate and real stock prices. Second, stock prices are an important determinant since cointegration fails if real stock prices are left out. Third, economic and financial reforms did affect the stability of demand for money in Nigeria over the period 1986:1-2012:4. Fourth, the long run income elasticity is not significantly different from unity with the inclusion of real stock prices. Finally, the results show structural breakpoints and they look to match clearly with corresponding critical financial, economic and political incidents.

JEL : E41, E44, C12

KEYWORDS: Money Demand, Structural Break, Cointegration, Nigeria

INTRODUCTION

Over the years, several studies have examined various related issues on money demand function (MDF) both in the developed and developing countries. The extensive and ever increasing stream of empirical research on MDF has to do with its importance in the design and implementation of an optimal monetary policy (Sriram 2001, Niyimbanira 2009, and Bahmani-Oskooee and Gelan 2009). For example, income elasticity of money demand gives us the long-term consistent rate of monetary expansion while interest elasticity of money demand helps in calculating the welfare cost of long-term inflation. Asides, the stability of MDF is fundamental to the choice of appropriate monetary policy actions in any economy. Though a wealth of studies has established a long term relationship between real balances and income, yet there remains uncertainty regarding long-run demand elasticity. This is particularly critical to developing countries like Nigeria that has undergone significant systemic changes since 1986. In addition, the role of stock market in determining money demand has not been assessed in the case of Nigeria in spite of the phenomenal growth of the stock market since 1986.

As pointed by Friedman (1988), stock prices have two kinds of impacts on money demand, a positive wealth effect and negative substitution effect. A wealth effect occurs in three different scenarios. First, a rise in stock prices leads to additional wealth which may be stored in money. Two, an increase in stock prices reflects an increase in the expected return from risky assets relative to safe asset. The resulting increase in relative risk may induce economic agents with given risk aversion /preference to hold larger amounts of safer assets such as money in their portfolio. Third, a higher level of stock prices may imply a rise in the volume of financial transactions, which may result in an increase in money demand to facilitate these transactions. On the other hand, the negative substitution effect of stock returns in money demand implies that as stock prices rise, the economic agents may preferably hold larger equities to other component of the portfolio, as the equities became more attractive or profitable (Thorton 1998, Choudhry 1996). Studies have documented both the negative and positive effects of stock prices on demand for

money for many developed and industrialized economies (Hsing, 2008, Baharumshah, *et al*. 2009, Serletis, 1993). However, for sub Saharan Africa, not many studies except Dube (2013) have attempted empirical verification of this issue in spite the various capital market reforms in the region. Hence, it is intriguing to investigate whether or not stock prices matter for demand for money function using Nigeria as a case study. Moreover, the issue of stability of the MDF, particularly in the developing economies, has come to the front burner following the plethora of financial reforms instituted since 1980s.

It is argued that the implementation of these reforms might have made MDF unstable thereby affecting the effectiveness of monetary policy in these countries. This issue is particularly important for Sub-Saharan Africa countries like Nigeria considering the fact that unstable money demand caused by financial reforms in the late 1970s actually induced many central banks in developed countries to switch from money targeting to the interest rate as monetary policy instrument. This is empirical issue that needs to be investigated. Therefore, the objective of this paper is threefold. One, ascertain the integrating properties of real money balances (broad and narrow), income, interest rate, exchange rate and stock prices in Nigeria. Second, examine the size and direction of the effect of stock prices on the demand for money. Three, determine the stability of the parameters of the long run money demand function by applying the unit root tests and cointegration tests with structural break. The rest of the paper is organized as follows: section 2 provides a brief summary of empirical studies on money demand and its stability. Section 3 contains the methodology and the data. In section 4, we present the empirical results and the final section contains the concluding remarks.

LITERATURE REVIEW

Studies abound on the demand for money and its stability. These studies include Mckinnon, et al. (1984), Arango and Nadiri (1981), Bahmani-Oskooee and Pourhaydrain (1990), Rao and Kumar (2009), Biscari et al. (2010), and Zuo and Park (2011) among others. However, for brevity we only provide a capsule summary of the empirical studies on the role of stock market on money demand and its stability. The study by Choudhry (1996) examined the role of stock market on money demand for US and Canada. The results show the stock market performance of these two countries has a significant role in the long run real M1 and M2 demand functions. In the same way, Thorton (1998) show that real stock prices have a significant and positive effect on the long run demand for narrow money balances in Germany. Biggs (2003) and Cassola and Morana (2004) provide evidence of significant long run elasticities with respect to the stock market price index but with opposite sign. Kontolemis (2002) and Bruggerman *et al*. (2003) conclude that stock prices do not matter for money demand in the long run, but may be useful in forecasting exercises. The study by Bruggeman *et al*. (2003) do not find stock market volatility as a significant factor explaining money demand relation. However, Carstensen (2006) reveal a significant long run impact of stock market volatility on money demand in the euro area. The study found that money demand model that incorporates stock market volatility and equity returns exhibits structural stability.

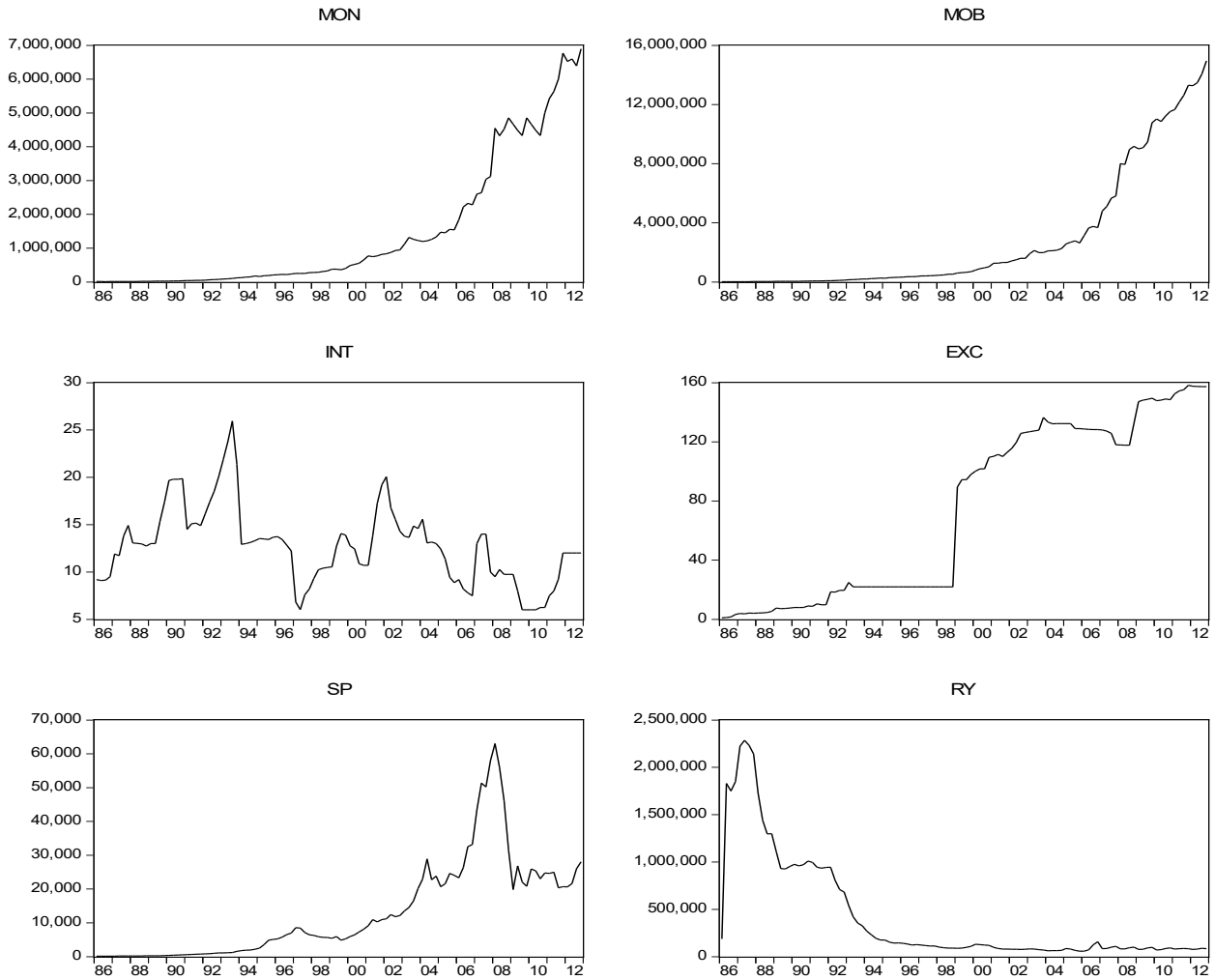
With reference to Nigeria, several studies have examined the stability of the money demand function. These include Arize (1984), Darrat (1986), Adejugbe (1988), Amadi, (1999), Nwaobi, 2002, Fielding (1994), Anoruo (2002), Akinlo (2005), Owoye and Onafowora (2007), Nwafor *et al* (2007), Ajewole (1989), Chukwu *et al*. (2010) and Kumar *et al.* (2013). In general, controversy remains in the literature on the stability or otherwise of the both the narrow and broad money. For example, study by Anoruo (2002) found broad money demand function was stable during the period 1986-2000. In the same way, Akinlo (2005) and Owoye and Onafowora (2007) found the broad money demand function was stable in Nigeria. However, the study by Kumar *et al*. (2013) found that the narrow money demand function was stable over the period 1960-2008. One major lacuna in all the existing empirical studies on money demand especially in Nigeria is that they failed to consider the role of stock prices. There is the need to address this gap in the literature.

DATA AND METHODOLOGY

Data

The study covers the period 1986:1 to 2012:4. Quarterly data for narrow money (MON) defined as currency plus demand deposits held by households and enterprises; broad money (MOB) defined as MON plus quasi money, real income (ry) defined as nominal GDP deflated by the consumer price index, exchange rate (exc) defined as units of domestic currency per unit of U.S. dollar and interest rate (int) were used. All quarterly series of the variables were obtained from Central Bank of Nigeria, Statistical Bulletin 2012 and 2013 editions. The plots of the variables are as shown in Figure 1.

Figure 1**:** The Plots of MON, MOB, Interest Rate, Exchange Rate, Stock Price and Real Income



This figure shows the trends of MON, MOB, interest rate, exchange rate, stock price and real income.

Model Specification

Most previous studies are based on the following general specification of the standard semi-logarithmic specification of the long-run money demand function:

$$\frac{M^d}{P} = f(Y_{t,}\, i) \tag{1}$$

where the demand for real balances $^{m}/_{p}$ is a function of the chosen scale variable (Y) to represent economic activity and the opportunity cost of holding money (i). M stands for selected monetary aggregates in nominal term and p for the price level. However, in this work, we specify the demand for money function for Nigeria to include exchange rate and stock prices as:

$$ln\left(\frac{m}{p}\right)_t = \emptyset_0 + \emptyset_1 ln(ry)_t + \emptyset_2 int_t + \emptyset_3 e_t + \emptyset_4 ln(sp)_t + \in_t \tag{2}$$

where M/P is the real money stock, ry is logged real income measured as real income $\left(^{GDP}/_{CPI}\right)$, int is the rate of interest, exc is the exchange rate and sp is real stock prices. $\in_t$ is the residual term, assumed to be a white noise process. Usually, real GDP (ry) represents the real income and, therefore, the transactions volume in the economy. The opportunity cost of holding money is proxied with nominal interest rate. In the same way, as contended by Friedman (1988), since stock prices measures the wealth effect, the stock prices should be measured in real rather than in nominal term.

The expectation is that the income elasticity of money demand be positive ($\emptyset_1 > 0$). More precisely, the quantity theory of money proposes a value of 1 for $\emptyset_1$, whereas the Baumol-Tobin model predicts a magnitude of 0.5 for $\emptyset_1$. The interest rate (int) is the rate of return at which economic agents prefer holding some alternative (financial or physical) assets as against money. Consequently, the anticipated sign for the semi-elasticity for interest rate is negative ($Ø_2 < 0$). The exchange rate is incorporated based on the literature on currency substitution, which suggests that portfolio shifts between domestic and foreign money can be captured by the exchange rate. The sign of the elasticity of exchange rate is ambiguous ($0 < Ø_3 < 0$) depending on the strengths of the substitution and income effects. There is a positive currency substitution effect when a stronger domestic currency (exchange rate appreciation) increases domestic money demand. On the contrary, a real exchange rate appreciation is associated with a negative shock to economic activity which could lower domestic money demand. Finally, the stock price is incorporated following the argument of Friedman (1988) as earlier discussed in the introduction. The expected sign is ambiguous ($0 < Ø_4 < 0$), depending on the relative strengths of the income and substitution effects. However, if $Ø_4 = 0$, it means that there is no role for stock market as a determinant of demand for money in Nigeria.

Method

The study adopts co-integration with structural break tests. However, before applying this approach, the integration properties of individual series need to be investigated to detect whether the series contains unit roots. To this end, the standard augmented Dickey Fuller (ADF) and Kwaiatkowski-Phillips-Schmidt-Shin (KPSS) (1992) are adopted. Specifically, we apply Dickey and Fuller's (1981) three model tests. The model follows the determining rule of Doldado *et al*. (1990) to establish the appropriate model. Moreover, in selecting the lag length, the paper adopts the modified Akaike Information Criterion (MAIC) suggested by NG and Perron (2001). This is done to avoid possible bias if the lag length is pre-determined without any rigorous determination approach. However, conventional ADF and KPSS tests may nevertheless be suspect, by not taking into account that the structural breaks could lead to a wrong decision when the null hypothesis is not rejected (Lee and Chien, 2008). As pointed out by Hendy (1996), a structural break relates to an intermittent shock with a permanent effect on the series. The obverse can occur where the break manifests at the beginning of the sample (Leybourne *et al*. 1998). Taking into consideration this possible shift in regime in the unit root tests, Zivot and Andrews (1992) developed a new category of tests that account for a structural break. The three models of the ZA tests are as expressed in equations 4-6.

Model A:

$$\Delta Y_t = {\mu_1}^A + {\gamma_1}^A t + {\mu_2}^A DU_t(\lambda) + \alpha^A Y_{t-1} + \sum_{j=1}^{k-1} \beta_j \Delta Y_{t-j} + \epsilon_t \quad (3)$$

Model B:

$$\Delta Y_t = {\mu_1}^B + {\gamma_1}^B t + {\gamma_2}^B t DT_t^*(\lambda) + \alpha^B Y_{t-1} + \sum_{j=1}^{k-1} \beta_j \Delta Y_{t-j} + \epsilon_t \quad (4)$$

Model C

$$\Delta y_t = {\mu_1}^C + {Y_1}^C t + {\mu_2}^C Du_t(\lambda) + {Y_2}^C DT_t^*(\lambda) + \alpha^C Y_{t-1} + \sum_{j=1}^{k-1} \beta_j \Delta Y_{t-j} + \epsilon_t \quad (5)$$

where $DU_t(\lambda)$ is a dummy variable which is equal to 1 and $DT_t^*(\lambda)$ = t – Tλ, 0 otherwise. In addition, λ = T_B/T, and T_B denotes a possible break point, where T is the sample size. The breakpoint is searched over the range of the sample (0.15T, 0.85T), and it can be estimated endogenously. Essentially, while model A permits a change in the level of the series, model B allows for a change in the shape of the trend of the series. Model C on the other hand, combines both changes in the level and the trend. In selecting the optimal length for ZA unit root test, we adopt the findings from the ADF tests.

EMPIRICAL RESULTS

Unit Root Test

The results of unit root tests using Augmented Dickey Fuller (ADF) and KPSS (1992) are as shown in Table 1. Overall, the tests show that all variables (including stock prices) possess a unit root in terms of level while they are all first difference stationary.

Table 1: The Results of Unit Root Tests

Test of Variables	**ADF Statistics**		**KPSS Statistics**	
Level	**No Trend**	**Trend**	**No Trend**	**Trend**
mob	0.023	-1.701	1.779	0.464
mon	-0.342	-1.829	1.822	0.408
ry	-2.866	-1.374	1.776	0.47
int	-2.384	-3.14	0.68	0.109
exc	-0.504	-2.108	2.113	0.198
sp	-2.042	-1.255	2.082	0.421
First Difference				
Δmob	-3.872	-4.005	0.314	0.071
Δmon	-4.596	-4.621	0.192	0.062
Δry	-4.178	-4.716	0.136	0.139
Δint	-5.557	-5.539	0.059	0.046
Δexc	-4.537	-4.514	0.089	0.084
Δsp	-3.903	-4.215	0.37	0.054

Critical values for ADF are -3.495, -2.495 and -2.582 (constant only at level); -3.496, -2.890 and -2.582 (constant only at 1st difference); and -4.050, -3.454 and -3.153 (constant and linear both level and 1st difference) at 1%, 5% and 10% level of confidence. The critical values for KPSS test are 0.739, 0.463 and 0.347 (constant at level); 0.216, 0.146 and 0.119 (constant and linear at 1st difference) at 1%, 5% and 10% respectively. The appropriate lag length for each variable based on MAIC given by Ng and Perron (2001) was 4.

Therefore, following the literature, we treated all the variables that appeared in the model 1 as *I*(1) process. However, to take into account the structural break that could lead to a wrong decision when the null hypothesis is not rejected; three models of ZA tests were conducted. The results of ZA test equations 3-5 are presented in Table 2 and depicted Figure 2 in Appendix 1. All series carry a unit root in the level and reject the null of non-stationarity in the first difference. This assists in ensuring that the *I*(1) type series for all series are considered. The ZA test results indicate that the breakpoints occur in 2001:Q3 for

broad money (MOB), 1993:Q3 for narrow money (MON), 1993:Q3 for real income, 2006:Q3 for stock prices, 1993:Q4 for interest rate and 2004:Q1 for exchange rate.

Table 2: The Results of ZA Unit-Root Tests with Structural Break

	Levels	Breakpoint	First Difference	Breakpoint
MOB	-2.778(C)	2001:Q3	-4.643(C)	2008:Q2
MON	-3.376(C)	1999:Q3	-7.029(C)	2007:Q4
ry	-3.679(C)	1999:Q3	-7.998(B)	1991:Q2
sp	-4.904(C)	2006:Q3	-5.401(B)	2006:Q4
int	-5.393(B)	1993:Q4	-7.598(B)	1994 :Q2
exc	-2.219(C)	2004:Q1	-10.14(B)	1999 :Q1

Critical values for models B and C are -4.80, -4.42 and 4.11 for model s B and C respectively, from Zivot and Andrews (1992). The characters within the parenthesis indicate the appropriate model according to the results from the ADF test.

However, it is possible for a series of data to contain more than one structural break. Consequently, we further employ the Bai and Perron (BP) (1998, 2003) approach to test for multiple structural breaks at unknown points in the sample. Table 3 shows the number of breaks selected by the BP sequential tests. For all the variables, the sequential procedure selects five breaks. First, we look at the double maximum tests of UD_{max} and WD_{max} that reject the null of no break vs. an unknown number of changes given the upper bound of five breaks.

Table 3: Bai and Perron's Test Results for Structural Breaks

Variable	U Dmax	W Dmax	Sup F(1/0)	Sup F(2/1)	Sup F(3/2)	Sup F(4/3)	Sup F(3/4)	No of Breaks
Mob	425.1**	736.5**	81.02**	69.94 **	72.559**	82.71**	85.00**	5
Mon	302.6**	487.3**	53.54**	47.86**	54.013**	60.51**	56.25**	5
Sp	257.4**	412.5**	18.62**	208.0**	257.4**	239.9**	143.5**	5
Ry	229.2**	394.1**	32.69**	28.71**	218.7**	229.2**	137.6**	5
Int	103.7**	177.2**	20.74**	16.56**	17.65**	18.79**	20.45**	5
Exc	2624.2**	3864.3**	176.5**	386.5**	481.5**	524.8**	355.5**	5

*Table 3 shows the Bai and Perron Test results for structural breaks. ** indicate significance at the 5% level. The upper bound M is set to be 5 and the trimming percentage is chosen to be at 15% in all cases.*

These tests show clearly that breaks exist. Furthermore, the Sup F(l+1/l) statistics (l = 1, 2, …,5) indicate the presence of five breakpoints. BP results for the location of structural breaks show that breakpoints occur in 1990:Q1, 1994:Q3, 1998Q4, 2003:Q4 and 2008:Q1 for real money balance narrowly defined (MON) while the breakpoints for real money balances broadly defined (MOB) are: 1990:Q3, 1994:Q3, 1999:Q1, 2003:Q3 and 2007:Q4. The breakpoints occur for real income (ry), in 1990:Q3, 1994:Q3, 1998:Q3, 2002:Q3 and 2006:Q3 for stock prices, 1990:Q1, 1994:Q1, 2000:Q1, 2004:Q2 and 2009:Q1 for the interest rate. With respect to exchange rate, the breakpoints are 1990:Q3, 1994:Q3, 1999:Q1, 2003:Q4 and 2009:Q1. Based on the results of ZA and BP tests for structural breakpoints, it is possible to summarize the following dates of structural breakpoints and find critical economic, political and financial incidents that can match the structural breaks of the series. The break in 1990 can be attributed to the introduction of several financial institutions and Schemes toward the end of 1989 and early 1990 to deepen and broaden the financial sector. Such institutions and Schemes included the Nigerian Deposit Insurance Corporation (NDIC) and the National Economic Reconstruction Fund (NERFUND) and Unit Trust Scheme. Moreover, the promulgation by the government, of the Banks and Other Financial Institution Decree (BOFID) No. 25 of 1991 served to enhance the control of Central Bank of Nigeria over the financial system. In particular, it enhanced the control of CBN over financial institutions that were traditionally outside the control of the Bank. Specifically in 1990, parastatals were compelled to withdraw their funds with the commercials and merchant banks and lodge same with the CBN. This led to a substantial deceleration in the growth of the money supply (M1) at 30.8 per cent as against 43.6 per cent in previous year. The break in 1994 could be attributed to the political crisis that followed the annulment of the 1993 election in the country. Government fiscal operation was characterized by huge fiscal deficits financed through credit from the Central Bank, which led to rapid expansion of broad money. The 1998

break might be attributed to the policy reversal that occurred during the period. Government reverted by to the policy of exchange rate pegging as against the floating system that was in operation after the deregulation in mid 1986. Moreover, Central Bank of Nigeria liquidated distressed 26 banks which led to major policy reforms in the banking sub sector. Government introduced tighter controls on bank lending to the private sector The date 2000 first quarter marked the re-introduction of government policy that compelled government parastatals and agencies to withdraw their funds from commercial banks which had earlier been cancelled. The breaks in 2002 and 2003 could be attributed to the various tight monetary and credit policies introduced by the government to prevent bank collapse due to low capital base, dominance of few banks, insolvency and illiquidity, overdependence on public deposits and foreign exchange trading, poor asset quality and low depositors' confidence.

The next date is 2004. There was consolidation of the banks in Nigeria from 89 to 24. In 2006, Government implemented the policy of recapitalization. The commercial banks were compelled to maintain a minimum paid up capital of ₦25 billion as the previous ₦2 billion. This policy forced many banks to merge and consolidate leaving only 25 in operations. The break in the 2007 could be explained in terms of the global financial crises that arose from US house bubble and the attendant increased default rates on mortgage. The Central Bank initiated several far reaching monetary and fiscal policies to mitigate the spillover effects of the crises on the domestic economy. In the year 2009, Central Bank introduced some measures to address the critical situation in the banking sector. The Audit test by the CBN in 2009 showed that 10 out of the 24 existing banks in the country had inadequate capital and liquidity to support their level of their current operations and future growth. Therefore, two (2) of the banks were asked to re-capitalized while eight (8) were adjudged to be in a distressed situations. The Executive Management of the 8 banks were sacked and Government injected ₦620 billion (about US $4.03 billion) in the eight bank to keep them operation.

Cointegration Test

Having established that the series are $I(1)$, the next issue is to determine whether there exists a long run equilibrium relationship of money demand function. To achieve this, we adopt the two well known statistics: the maximum eigen-value (λ-max) and the trace statistics proposed by Johansen and Juselius (1990). Some points are worth mentioning before presenting the results. First, we determine the number of lags (k) by using the Akaike Information Criterion (AIC) and Schwarz Bayesian Criterion (SBC). They both suggest four lags for the VAR. Two, the estimation procedure assumes unrestricted intercepts with no trend in the VAR model. The results in Tables 4a and 4b present the trace statistics and λ-max statistics for real money balances broadly and narrowly defined respectively.

Table 4a: Johansen Test for Co-Integration

Panel A:		**Variables**	**Mob, ry**	**sp, int, exc**	
H_0	H_1	λ - Max	CV (0.05)	Trace test	C.V. (0.05)
r = 0	r = 1	34.52	33.88	77.99	69.82
r ≤ 1	r = 2	20.80	27.58	43.47	47.86
r ≤ 2	r = 3	10.13	21.13	22.67	29.80
r ≤ 3	r = 4	6.840	14.26	12.54	15.49
r ≤ 4	r = 5	3.390	3.840	3.390	3.840
Panel B					
Estimated Co-Integration Vector Normalized on MON					
lnmon	lnry	lnsp	int	exc	
-1.0000	1.611	0.953	-0.119	0.011	
	(0.249)***	(0.206)***	(0.029)***	(0.003)***	

*Lag length selected by AIC is four. *** and ** denote significant at 1% and 5% level respectively. λ-max and trace test are Johansen's maximum eigen value statistics for testing co-integration. Critical values (C.V.) are from Osterwald –Lenum (1992). The values in parenthesis are the error terms*

Table 4b: Johansen Test for Co-Integration

Panel A:		**Variables**	**Mob, ry,**	**sp, int, exc**	
H_0	H_1	λ - Max	CV (0.05)	Trace test	C.V. (0.05)
r = 0	r = 1	35.30	33.88	81.33	69.82
r ≤ 1	r = 2	22.22	27.58	46.03	47.86
r ≤ 2	r = 3	11.38	21.13	23.81	29.80
r ≤ 3	r = 4	6.460	14.26	12.43	15.49
r ≤ 4	r = 5	3.590	3.840	3.590	3.840
Panel B: Estimated Co-Integration Vector Normalized on MOB					
lnmob	**Lnry**	**lnsp**	**int**	**exc**	
-1.0000	1.342	0.783	-0.108	0.010	
	(0.219)***	(0.179)***	(0.026)***	(0.003)***	

*Var Lag selected by AIC and used is shown. ***and ** denote significant at 1% and 5% level respectively. λ-max and trace test are Johansen's maximum eigen value and trace eigen value statistics for testing co-integration. Critical values (C.V) are from Osterwald –Lenum (1992). The values in parenthesis are the error terms*

Both test statistics suggest that there exist one co-integrating vector, which implies that there exists a long run relationship amongst the variables. We normalize the co-integrating vector with respect to the real money balances (MOB and MON) and the co-integrating relations are as shown in panel B of Tables 4a and 4b respectively. The exclusion tests suggested by Johansen and Juselius (1990, p 194) accept the null hypothesis that all the variables enter into the long term relationship. All the variables are statistically significant at 1% level. The results show that income and interest rate carry their expected positive and negative signs respectively. The exchange rate and stock prices both have positive sign supporting the wealth effect in the literature. The standard hypothesis in demand for money research is whether MDF is homogenous to degree one with respect to price level. To verify this, we regressed MOB and MON respectively on a constant, real income, stock prices, interest rate, exchange rate and price level (log of consumer price index). A Wald test that the estimated long-run coefficient on the price level is unity yielded a value of χ^2 = 2.399 (p = 1.220) and χ^2 = 0.648 (p= 0.4207) for narrow (MON) and broad money (MOB), which are larger than the 5% level of significance. Consequently, we could not reject the null hypothesis of unitary price elasticity at reasonable significance levels.

The finding that the demand for money is homogenous to degree 1 with respect to price level means that either narrow or broad monetary measures could be used to determine the long run effect of the monetary policy (Hafer and Kutan 1994, Bahamrunshah *et al*. 2009). The results show that income elasticity for narrow money is unity (χ^2 = 30.94; p = 0.0000) likewise for broad money aggregate (χ^2 = 16.208; p = 0.0001). It is very important to note that the magnitude of the scale variable (1.611) for (broad money) and (1.342) for narrow money aggregates are still considered high compared with what was reported for developed and some Asia countries (see Mark and Sul 2003, Harb, 2004; Kumar et al 2010; Hafer and Jansen, 1991; Drake and Chrystal 1994; Haug and Lucas 1996 and Lim 1993). This simply suggests that the availability of alternative financial assets as money substitutes in household portfolio is still low in the country. However, these magnitude of scale variable for both narrow and broad money aggregates are quite close to values obtained by earlier studies for some newly industrialized economies such as China, Malaysia (see Deng and Liu 1999; Chen 1997; Arize and Malindretos 2000). There is therefore the need for more reforms in the financial sector to bring about greater competition and development of alternative financial assets that could serve as alternatives to money. We robust check the results by estimating the model with restricted trend. The results show no cointegration between money and its determinants. The coefficient still remained significant but the scale variable dropped to 1.171 and 1.039 for broad and narrow money respectively. In addition, the trend variable was significant.

Table 5 shows the sort-run dynamics of the error-correction model. The question that arises is how robust is the results to the exclusion of stock prices? The exclusion of stock prices leads to no co-integration between money and its determinants (results not reported for space consideration). This is in line with the findings of Wu, *et al*. (2005) and Baharumshah, *et al.* (2009). However, the coefficient of income declines to less than one for both real M1 (0.6014) and real M2 (0.6923) but remain significant with the exclusion of stock price. The coefficient of interest rate increases for both real M1 (0.1689) and real M2 (0.1891).

Table 5: Results for Error Correction Model

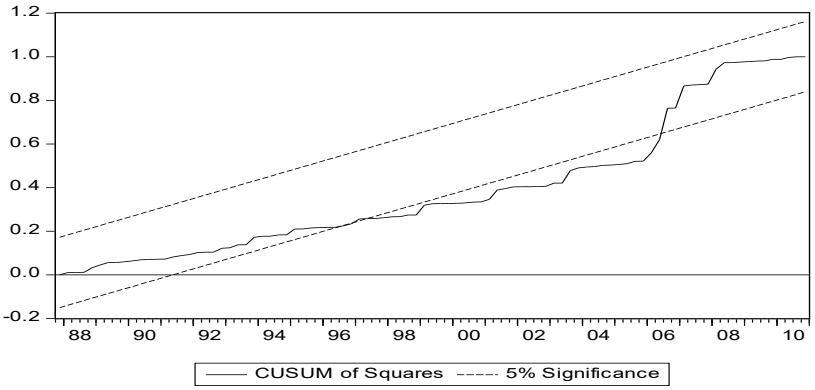
Variables	$\Delta log(MOB)$	$\Delta log(MON)$
$\Delta log(BM)_{-1}$	0.062	
	(0.56)	
$\Delta log(BM)_{-2}$	0.0113	
	(0.11)	
$\Delta log(BM)_{-3}$	-0.043	
	(-0.39)	
$\Delta log(BM)_{-4}$	0.102	
	(1.03)	
$\Delta log(NM)_{-1}$		0.061
		(0.61)
$\Delta log(NM)_{-2}$		-0.014
		(-0.14)
$\Delta log(NM)_{-3}$		-0.049
		(-0.49)
$\Delta log(NM)_{-4}$		0.1774
		(2.10)**
$\Delta log(RY)$	0.403	0.479
	(5.53)***	(6.14)***
$\Delta log(RY)_{-1}$	0.087	0.0002
	(0.97)	(0.00)
$\Delta log(RY)_{-2}$	-0.112	-0.068
	(-1.14)	(-0.65)
$\Delta log(RY_{-3}$	-0.049	-0.157
	(-0.52)	(-1.55)*
$\Delta log(RY)_{-4}$	-0.059	-0.077
	(-1.73)*	(-1.98)**
$\Delta log(SP)$	0.148	0.135
	(1.89)*	(1.59)*
$\Delta log(SP)_1$	-0.24	-0.051
	(-0.31)	(-0.62)
$\Delta log(SP)_2$	-0.072	-0.038
	(-0.91)	(-0.44)
$\Delta log(SP)_3$	-0.016	0.015
	(-0.20)	(0.18)
$\Delta log(SP)_4$	0.089	0.155
	(1.21)	(1.95)**
$\Delta log(INT)$	-0.001	-0.001
	(-0.19)	(-0.16)
$\Delta log(INT)_1$	0.005	0.001
	(0.80)	(0.09)
$\Delta log(INT)_2$	0.0001	0.003
	(0.01)	(0.46)
$\Delta log(INT)_3$	-0.0004	-0.005
	(-0.08)	(-0.82)
$\Delta log(INT)_4$	0.018	0.023
	(3.49)***	(4.00)***
$\Delta log(EXC)$	0.002	0.002
	(1.38)	(1.66)*
$\Delta log(EXC)_1$	-0.002	-0.002
	(-0.70)	(-1.29)
$\Delta log(EXC)_2$	-0.0004	-0.001
	(-0.40)	(-0.78)
$\Delta log(EXC)_3$	0.001	0.001
	(0.70)	(1.03)
$\Delta log(EXC)_4$	0.001	0.001
	(0.05)	(0.64)
Constant	-0.903	-0.957
	(-2.21)**	(-255)**
$EC)_{t-1}$	-0.052	-0.071
	(-2.31)**	(-2.63)**
R^2	0.54	0.55
Norm (2)	7.663[0.021]	3.636[0.162]
LM Test	0.211[0.900]	1.441[0.487]
ARCH(6)	4.864[0.561]	9.943[0.127]

*The number in parenthesis ( ) are t values while those in brackets [ ] are p- values . ***, ** and * denote significance at 1%, 5% and 10% level respectively. ARCH (m) is a m^{th} order test for autoregressive conditional heteroskedasticity; Norm (2) test indicates whether the residuals are normally distributed and LM test for autocorrelation*

Two, by replacing stock price with stock market capitalization, the results show a single unique co-integrating vector between real MON, output, interest rate, market capitalization and exchange rate. However, no co-integration between broad money aggregates (MOB) and its determinants. In both cases, all the coefficients are significant at 1% level. The coefficient of scale variable declines to 0.8714 for real MOB and 0.7612 for real MON. This might suggest that portfolio holders perceive stocks as a major substitute to money. Given co-integration as revealed in tables 4a and 4b, we investigate short run dynamics that include information on both long run and short run parameters, where the former is captured through the error-correction term (ECT). Tables 5 columns 1 & 2 summarize the short run dynamics of the error correction model (ECM) for broad and narrow money respectively. The coefficient of the error correction term (ECT_{t-1}) for both narrow and broad monetary aggregates (MON) and (MOB) carry the correction sign (negative and relatively small). The results show that the adjustment to an exogenous shock is rather slow.

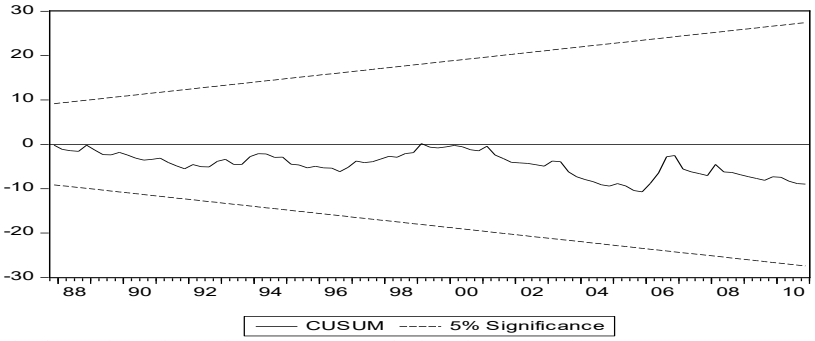
The coefficient of the ECT_{t-1} (0.054) reveals that approximately 5% of the previous quarter's discrepancy between the actual and equilibrium value of broad aggregates is corrected each quarter. The corresponding percentage for narrow money aggregate is 0.07. Essentially, co-integration amongst broad and narrow money aggregates and their determinants (stock prices inclusive) can be confirmed by the significance of the lagged error-correction term. Real income and stock prices current value have significant short term effect on money demand (narrow and broad). However, domestic interest rates, which appear to be an important variable for estimating long-run co-integrating vector, has a non significant short term effect on money demand. Exchange rate current value has positive short-term effect on money demand but significant only at 10% for narrow money demand. We proceed further to ascertain the stability of the long run coefficients that are used to form the error correction term in conjunction with the short run dynamics. For this purpose, we use the CUSUM and CUSUMSQ tests proposed by Brown, Dublin and Evans (1975).The tests are applied to the residuals of both narrow and broad money. The CUSUM test is based on cumulative sum of the recursive residuals. This option plots the cumulative sum together with the 5% critical lines. The test indicates parameter instability if the cumulative sum goes outside the area between the two critical lines. The same procedure is adopted to carry out the CUSUMQ, which is based on the squared recursive residuals. The graphical presentation of these two tests for the narrow and broad money is as shown in Figures 3(a-d). The CUSUM tests show evidence of stability of two money aggregates (Figures. 3b & 3c). However, the cumulative sum of squares shows evidence of parameter instability for both narrow and broad money (Figures (3a & 3d).

Figure 3a: CUSUM of Squares for Broad Money (MOB)



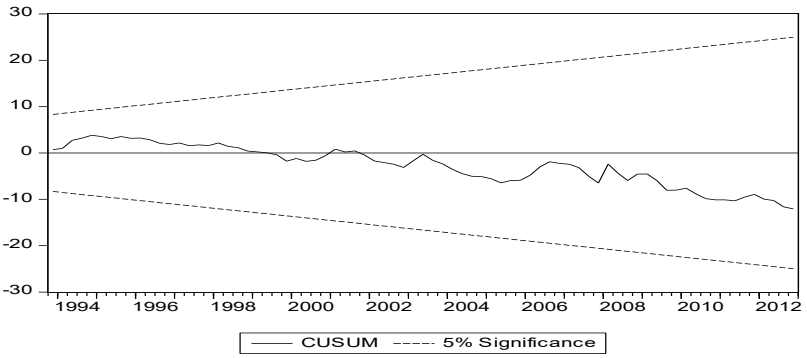
This figure shows the result of CUSUM of Squares for Broad Money (MOB)

Figure 3b: CUSUM Test for Broad Money (MOB)



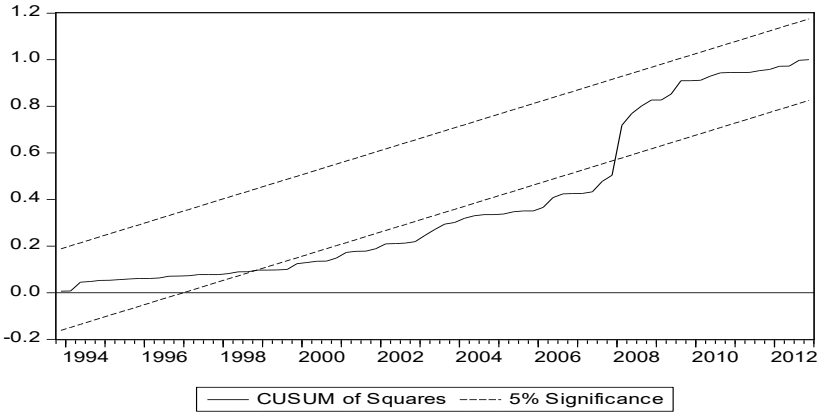
this figure shows the result OF CUSUM test for broad money (mob)

Figure 3c: CUSUM test for narrow Money (MON)



This figure shows the result of Cusum test for Narrow Money (MON)

Fig 3d: Cusum of Squares Test for Broad Money (MOB)



This figure shows the result of Cusum of Squares test for broad money (MOB)

Structural Break Test

However, the standard co-integration test may be inappropriate in the presence of structural breaks. Therefore, it is necessary to check the co-integration relationship for structural breaks. To accomplish this, the study adopts the Gregory and Hansen (1996) approach and test for co-integration between variables in the models with regime shifts. The Gregory and Hansen test is based on the notion of regime change and is a generalization of the usual residual based co-integration test. The GH test has a null

hypothesis of no co-integration and its alternative hypothesis suggest co-integration with one structural break. The co-integration between variables exists when the null hypothesis suggests co-integration with one structural break. Gregory and Hansen create three models as follows:

Level shift (C)

$$y_t = \delta_0 + \delta_1 \emptyset_t(\tau) + \alpha(X_t) + e_t \quad (6)$$

Level shift with trend

$$(C/T) \qquad y_t = \delta_0 + \delta_1 \emptyset_t(\tau) + \beta(t) + \alpha(X_t) + e_t \quad (7)$$

Regime shift with slope vector shift

$$(C/S) \quad y_t = \delta_0 + \delta_1 \emptyset_t(\tau) + \alpha(X_t) + \beta X_t \emptyset_t(\tau) + e_t \quad (8)$$

Where the structural shift in each equation is shown by a dummy variable ∅ and defined as:

$$\emptyset_1 = \begin{cases} 0, if\ t \leq T_\tau \\ 1, if\ t > T_\tau \end{cases}$$

Here $\tau = T_B/T$ and TB represents a possible breakpoint. GH have developed versions of the co-integration ADF tests of Engle and Granger (1987), as well as the Z_t and Z_α tests of Philips-Quliaris (1990), whereby all of them are modified according to the alternative considered. Taking into account that the date of the change is unknown, GH compute the values of ADF* = $\inf_{\lambda \epsilon j}$ ADF, Z_t* = $\inf_{\lambda \epsilon j} Z_t$ and Z_α* = $\inf_{\lambda \epsilon j} Z_\alpha$. This model is estimated recursively allowing the breakpoint τ to vary such that $[0.15T \leq \tau \leq 0.85T]$. The results of the GH test are as presented in table 6. The results of both narrow (MON) and broad (MOB) money aggregates show clear evidence of cointegration even when we allow for structural break. Both narrow and broad money demand function reveal a structural break in three models. This test, therefore, affirms the importance of a structural change in the co-integration vector and thus the need to consider it in the specification of the money demand function. This by implication means demand for money specification that envelopes the changing economic and financial incidents raise some critical questions on the long run relationship among the series (Lee and Chien, 2008). According to the three tests statistic criteria in the broad money aggregate function, the structural break years and quarters estimated are mainly 2006:Q2, Q3 and Q4, and 2008:Q1&Q2. In the case of narrow money aggregate function, the structural break years and quarters are namely; 2006:Q1, Q2 and 2007:Q1.

The structural break year 2006 quarters one to four can be attributed to the financial reforms implemented in the country. In an attempt to make the banking sector solid and sound, the Central Bank of Nigeria embarked on the policy of bank recapitalization. Banks were required to increase their minimum paid-up capital from ₦2 billion to ₦25 billion. The programme which started on January 1, 2006 led to the emergence of 25 banks from the existing 89 banks. Fourteen of the existing 89 banks that could not merge were liquidated. The ripple effects of this exercise made the Central bank to introduce many far reaching policies to stem the negative consequences of the distress in the banking sub sector on the economy. The breaks in 2007 and 2008 could be attributed to the global financial crises that started in 2007. The Central Bank of Nigeria instituted several exchange and interest rate policies to mitigate the effects of the economic meltdown in the domestic economy. These policies no doubt had significant effect on the demand for money.

Table 6: Gregory and Hansen (1996) Tests for Regime Shifts

	MOB		**MON**	
Model	Test Statistics	Breakpoint	Test Statistics	Breakpoint
ADF*				
C	-6.366**	2008 :Q1	-6.911**	2007 :Q1
C/T	-6.470**	2008 :Q1	-6.757**	2007 :Q1
C/S	-6.396**	2006 :Q3	-6.686**	2007 :Q1
Zt*				
C	-6.396**	2008 :Q1	-7.008**	2006 :Q1
C/T	-6.589**	2006 :Q4	-7.159**	2006 :Q1
C/S	-6.830**	2006 :Q2	-6.193**	2006 :Q2
Z*				
C	-60.66**	2008 :Q2	-67.82**	2006 :Q1
C/T	-62.52**	2006:Q4	-68.79**	2006 :Q1
C/S	-64.89**	2006:Q2	-56.65**	2007 :Q1

*Table 6 shows the results of Gregory and Hansen (1996) tests for regime shifts. :** significant at 5% level. The critical values are from Table 1 of GH (1996).*

Essentially, these tests provide evidence to support the fact that structural change is critical in co-integration vector. Consequently, it is necessary that it should be accounted for in the specification of money demand function. This finding is quite interesting as the endogenous estimation yields structure breaks that correspond to recognizable happenings both (financial and economic) in the economy. This, by implication, means that within the context of money demand, agents in the economy i.e. households and the government may respond differently when the economy is in a different regime. In short, money demand specification that does not take into consideration the various economic and financial events would cast doubt on the real money aggregate and its determinants.

CONCLUSIONS

This paper examines the role of stocks in the money function and tests for the stability of the function taking cognizance of the changing economic and financial incidents. The paper adopts co-integration approach that accounts for the possibility structural breaks with unknown timing. The results of the analysis show that stock prices have a significant substitute positive (effect) on long run narrow and broad money demand. The omission of stock prices could lead to serious misspecifications in the money demand function in both the short-and-long-run. The result shows that the demand for money is homogenous to degree 1 with respect to price level. The long-run demand for real money balances is negatively affected by the own rate of return for money. Exchange rate, however, has positive effect on the demand for money balances. The result of short run analysis shows that the adjustment process to an exogenous shock is rather slow given error correction coefficients of 0.054 and 0.07 for broad and narrow money aggregates respectively. The test, which allows for possibility of regime shifts, suggests a lack of stability in the demand for money given the data set from 1986:Q1 and 2012:Q4.

The ZA and BP tests revealed several structural break points and critical economic and financial incidents for matching with these break points.What policy conclusions can we draw from our empirical results? One, the results suggest that monetary policy aimed at stabilizing the domestic economy can generate only uncertainty if the effects of stock prices are not taken into consideration for the execution of monetary policy (Baharumshah, *et al*. 2009). Moreover, the finding that long run unitary income elasticity exists simply suggests that Friedman's rule is optimum in the case of Nigeria. This implies that money supply should grow at the same rate as output in order to attain the goal of price stability. The finding of structural breaks implies that while analyzing the demand for function of Nigeria, it is important to incorporate a structure change into the question. Finally, in both narrow and broad monetary aggregates, monetary income elasticity is far greater than interest rate elasticity. The policy inference from this is that in determining monetary policy, Central Bank of Nigeria should interrupt the quantity of money through real income as against changing the interest rate to be more efficient. The major limitation of this work is

that several variables including housing and land prices to model the wealth effect in money demand equation have not been incorporated. This is because data on many of these variables are not easily available in developing countries like Nigeria. There is, therefore, the need to source information on these variables in order to examine their roles in the money demand function in Nigeria.

Appendix 1

Figure 2a: Plot of Zivot-Andrews Unit Root Tests for BOM

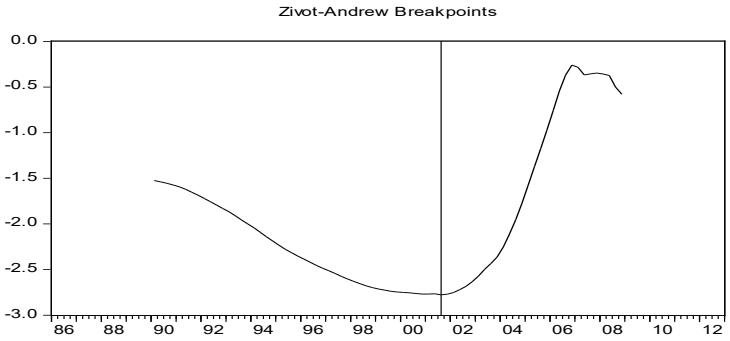


Figure 2b: Plot of Zivot-Andrews Unit Root Tests for MON

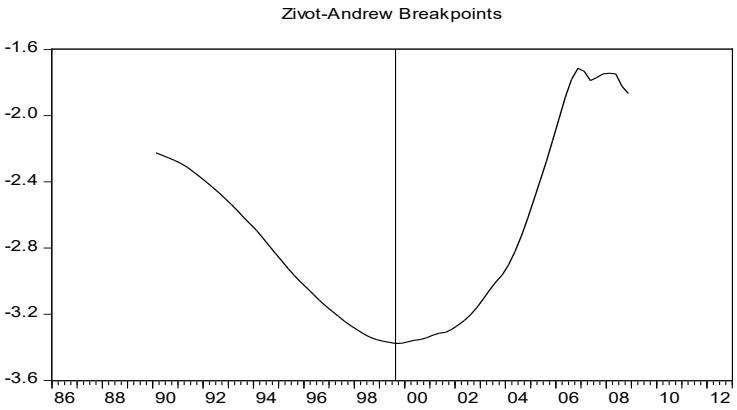


Figure 2c: Plot of Zivot-Andrews Unit Root Tests for SP

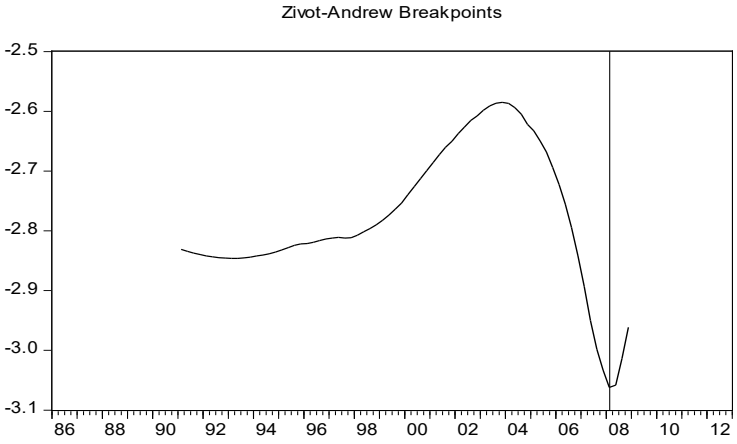


Figure2d: Plot of Zivot-Andrews Unit Root Tests for INT

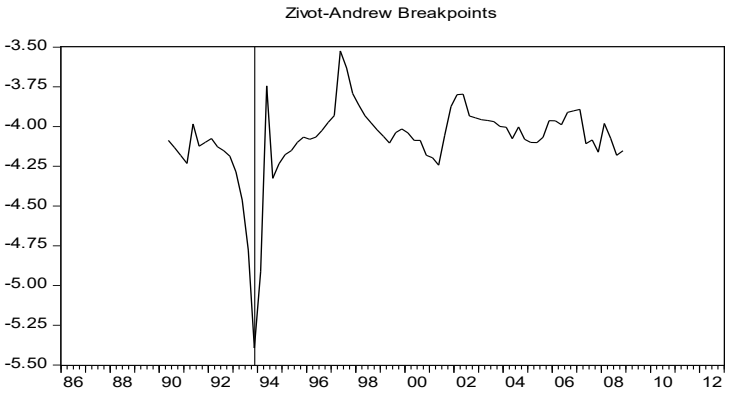


Figure2e: Plot of Zivot-Andrews Unit Root Tests for RY

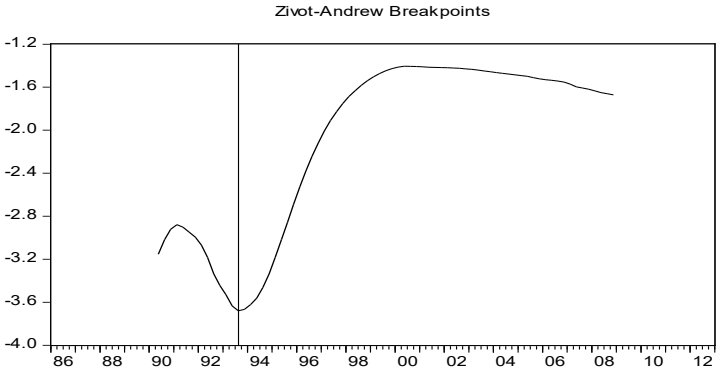


Figure 2f: Plot of Zivot-Andrews Unit Root Tests for EXC

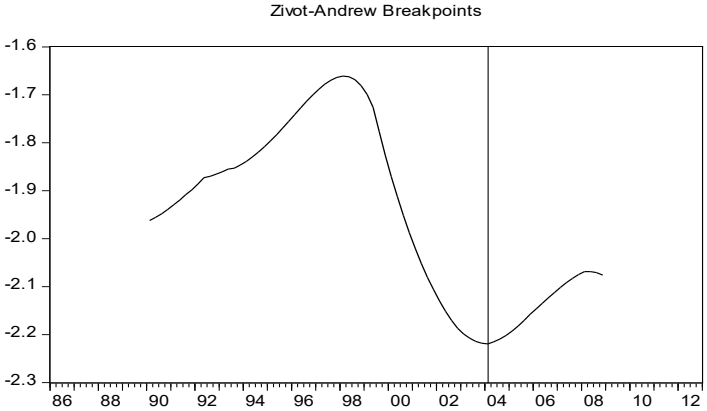


Fig. 2(a-f) Plots of Zivot and Andrews (1992) unit root tests for the variables

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