A VECTOR ERROR CORRECTION AND NONNESTED MODELLING OF MONEY DEMAND FUNCTION IN NIGERIA

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ABSTRACT

THIS PAPER EXAMINES THE STABILITY OF THE DEMAND FOR MONEY IN NIGERIA . WITH RELATIVELY SIMPLE MODEL SPECIFYING A VECTOR VALUED AUTOREGRESSIVE PROCESS(VAR), THE HYPOTHESIS OF THE EXISTENCE OF COINTEGRATION VECTORS IS FORMULATED AS THE HYPOTHESIS OF REDUCED RANK OF THE LONGRUN IMPACT MATRIX. THIS ENABLED US TO DERIVE ESTIMATES AND TEST STATISTICS FOR THE HYPOTHESIS OF A GIVEN NUMBER OF COINTEGRATION VECTORS . THE MONEY DEMAND FUNCTION WAS FOUND TO BE STABLE AND EVIDENCE GATHERED FROM THE NONNESTED TESTS SUGGEST THAT INCOME IS THE MORE APPROPRIATE SCALE VARIABLE IN THE ESTIMATION OF MONEY DEMAND FUNCTION IN NIGERIA.

JEL CLASSIFICATION: E41, C22, C32, C50

KEY WORDS: VECTOR ERROR CORRECTION MODE; COINTEGRATION; MONEY DEMAND; CONSUMPTION; NON-NESTED MODELS

1.0 INTRODUCTION

In most macroeconomic theories, the relation between demand for money balances and its determinants is a fundamental building block. And yet most macroeconomic models, whether theoretical or econometric, generally ignore the rich institutional detail of the financial sector and attempt to capture financial factors via the demand and supply of money. Furthermore, the demand for money is a critical component in the formulation of monetary policy and a stable function for money has long been perceived as a prerequisite for the use of monetary aggregates in the conduct of policy. This has therefore led to the extensive empirical scrutiny of demand for money function in many countries.

In the 1950s and the 1960s, theoretical microeconomic research on demand for money within a Keynesian framework stresses the constituent motives for holding money; while the macroeconomic analysis seemed to support Friedman's hypothesis that the demand for real money balances was a stable function of income and the interest rate (see Tobin, 1958; Friedman, 1956; Laidler, 1985; and Adams, 1992). However, following the oil-shock of 1973, many hitherto empirically robust relationships broke down with existing models persistently over-predicting actual real money holding (Goldfeld, 1976; Goldfeld and Sichel, 1990). Initial responses to this failure of aggregate demand for money functions focused on possible omitted-variable biases arising as a result of innovations and institutional change in financial systems. Attention was paid to the specification of domestic interest-rate effects and the consideration of the impact of currency substitution effects (Hendry, 1985 and Adam, 1991). Yet, this model re-specification proved to be a necessary but a sufficient condition. The issue of out of equilibrium behavior was broadly represented by the series are non-stationary. The presence of spurious regression results is high in view of the trending nation in the levels of the variables and they may fail adequately to capture the structural features among microeconomic variables.

These facts characterized most of the empirical money demand models of the Nigeria economy (see Tomori, 1972; Ojo, 1974; Odama, 1974; Teriba,1974; Ajayi, 1974, Ojo, 1974; Iyoha, 1976; Akinnifesi and Phillips, 1978; Fakiyesi, 1980; Darrat, 1986; Asogu and Mordi, 1987; Afolabi, 1987; Adejugbe, 1988; Audu, 1988; Ajewole, 1989; World Bank, 1991; Oresotu and Mordi, 1992; Essien, et al 1996). The adoption of partial adjustment framework may not appropriate for the real world dynamic economic modeling since the lag structure imposed by such framework may be different from the lag structure that characterizes the phenomenon being studied.

The appropriate specification of the relationship between the long run theory and the short-run dynamics has dominated much of the time series economic research in the 1980s and represents the principal response to the collapse of many of the aggregate macro-economic relationships in the 1970s (Davidson et al, 1978). Thus, the econometrics of dynamic specification have led to important revisions to the modelling of macro-economic relationships in recent years, including money demand functions (Engle and Granger, 1987; Johansen and Juselius, 1990). In line with this development, most of the recent applications for the Nigerian economy include studies by Teriba (1992), Nwaobi (1993a) Nwaobi (1993b) and Teriba (1994). And yet a more recent development is concerned with the appropriate scale variable to be utilised in the demand for money relationship. Most theoretical considerations suggest as a more appropriate scale variable in the demand for money, consumer expenditure rather than the traditional income variable (see Mankin and Summers, 1986; Arestis and Demetriades, 1991; and Elyasiani and Zadeh, 1995). This paper therefore sets out to examine the stability of the demand for money in Nigeria by concentrating on these developments.

In seeking to construct an improved model, we formulate an equation that integrates long-run properties with short-run dynamics, based on the recent merging of the theories of error correction and cointegration. The resulting model is critically evaluated. The sequence involves a natural progression from model discovery to model evaluation through replication and testing, and then via new conjectures back to discovery, seeking models that account for previous findings and explain additional phenomena. It is thus the objective of this paper to achieve this last goal for Nigeria's money demand models over the period 1960-1995. An additional objective is to exposit an econometric framework that makes precise the notion of an improved model, explains the construct of accounting for previous findings (denoted encompassing), and delineates the criteria for model evaluation. This will then enable us to clarify the concepts of encompassing and cointegration in the context of current important economics debate. Section two looks at the theoretical and data considerations. Section three analyses the cointegrating and dynamic relationships. The performance of the two estimated dynamic demand for money function is compared through non-nested procedures in section four. Section five summarises the argument and concludes the paper.

2.0 <u>THEORETICAL AND DATA CONSIDERATIONS</u>

The quantity theory of money started with the identity:

$$MV = PT \tag{2.1}$$

Where M is the quantity of money, V is the velocity of circulation, P is the price level, and T is the volume of transactions. Keynes further modified this identity by distinguishing three motives for holding money-transactions, precautionary, and speculative motives. Latter developments amplified on these Keynesian motives in various ways (see Goldfeld and Sichel, 1990). Baumol (1952) and Tobin (1956) both applied inventory-theoretic considerations to the transactions motive. In the simplest of these models, individuals are paid (in bonds) an amount Y at the beginning of a period and spend this amount uniformly over this period. This leads to the so-called square root law with average money holding given by

$$M = (2bY/r)^{½} \tag{2.2}$$

Where r is the interest rate on bonds and b is the brokerage charge or fixed transactions cost for converting bonds into cash. In another application of the inventory theory, Millers and Orr (1966) considered two assets and transactions cost, which are fixed per transaction. Given a lower bond below which money balances cannot drop (normalised to zero), the optimal policy consists of an upper bound, h, and a return level, Z. Whenever, money balances reach the lower bond, Z dollars of bonds are converted to cash; whenever the upper bound is reached, h - Z dollars of cash are converted to bonds. Minimising the sum of expected per-day transactions and opportunity costs yield the optimal return level:

$$Z^* = [(3b/4r)\sigma^2]^{\frac{1}{3}} \tag{2.3}$$

Where σ^2 is the daily variance of changes in cash balances (σ^2=M^2t).

In the two-asset version of the portfolio approach (reformulated Keynes speculative motive), Tobin (1958) showed that the individual wealth holder allocates his portfolio between money, treated as a riskless asset, and an asset with an uncertain rate of return. Under the assumption of expected utility maximization the optical portfolio mix can be shown to depend on wealth and on the properties of the utility function and the distribution function for the return on the risk asset. In the general multi-asset case however, the demand functions for each asset in the portfolio, including money depend on all the expected return and on the variances and covariances of these returns. From a theoretical point of view, the analysis yields negative interest elasticity for the demand for money, providing another rationalisation of keynes liquidity preference hypothesis. Friedman's (1956) restatement of the quality theory parallel Tobin's portfolio approach in regarding the primary role of money as a form of wealth. He treats money as an asset yielding a flow of services to the holder. Wealth, both human and nonhuman, is thus one of the major determinants of money demand. However, his approach side-steps the explicit role of money in the transaction process and also ignores problems of uncertainty.

In a partial attempt in resolving this problem, Ando and Shell (1975) consider a world with three assets, one risky and two, money and saving deposits, with certain normal returns, r_s and r_m. They also treat the price level as uncertain and view individuals as maximising expected utility is given by $U(C_1C_2)$ and C_i is the consumption in the i^{th} period. The role of money in the transaction process is captured in a real transactions cost function, $T(M,C_1)$, where holding higher money balances reduces transactions costs and thus, other things equal, raise C_2. They then assume that C_1 is determined independently of portfolio choice and show that the appropriate marginal condition for maximising expected utility is given by

$$r_s - r_m = T_m (M,C_1) \tag{2.4}$$

(2.4) can be inverted to give money demand as a function of C_1, (r_s-r_m), and the parameters of T. In other words, the demand for money is seen to be independent of the rate of return on the risky asset and of the expected price level and of wealth as well.

Similarly, McCallum and Goodfriend(1987) in a three-asset world consisting of money, bonds and capital; consider the case of certainty, and start an intertemporal household utility function of the form: $U(C_t L_t) + \beta U(C_{t+1}, L_{t+1}) + \beta^2 U(C_{t+2}, L_{t+2}) + \ldots$,

Where C_t and L_t are consumption and leisure. The household has a production technology as well as initial real stocks of money (M_{t-1}), bonds (b_{t-1}), and capital (k_{t-1}). Here, the role of money is captured by a "shopping time" function $S_t = \psi(C_t,M_t)$, where shopping time, S_t, subtracts from leisure. They then show that maximising utility results is a demand for money which can be written as

$$Mt = f(M_{t-1}, K_{t-1}, b_{t-1}, Rt, R_t, R_{t+1} \ldots, \prod_t, \prod_{t+1} ..) \tag{2.5}$$

Where R_t and $\prod_i$ are the nominal interest rate and the inflation rate, respectively, and where variables dated after t are anticipated values. After some manipulation, they further showed that (2.5) can be transformed to

$$M_t = g(C_t, R_t) \tag{2.6}$$

(2.6) results from the fact the structure of the McCallum-Goodfriend model is such that the use of the choice variable, C_t, in (2.6) allows one to eliminate everything but R_t from (2.5). Of particular interest is the role of initial wealth, which appears in (2.5) via M_{t-1},b_{t-1}, and K_{t-1} but which has not in (2.6). Indeed, the bulk of empirical work on money demand has been motivated by one or more of these theories.

In Nigeria, money held with the commercial banking system exists as the principal financial asset. Alternative financial asset are scarce, while equity and bond markets are thin. Also, borrowing and deposit interest rates have traditionally been administered by the monetary authorities and for much of the period since 1960s, domestic interest rates have been monitored at low or negative levels in real terms. Given this background, we define the vector of variables of interest in determining the demand for money in Nigeria, X as

$$X = [LM1, RLSC, LINT, LP] \tag{2.7}$$

$$= \quad [LM1, RLGDP, LP, LINT] \quad (2.8)$$
$$= \quad [LM1, RLC, LINT, LP] \quad (2.9)$$

where LM1 is Natural Logarithm of nominal money balances; RLSC is the natural logarithm of real scale variable; LINT is the natural logarithm of interest rate; LP is the natural logarithm of prices; RLGDP is the natural logarithm of income variable; and RLC is the natural logarithm of real consumption variable. A number of important issues arise when considering the appropriate data to be used as proxies for the variables of the model. Here, we adopt the money aggregate report by the international momentary fund financial statistics yearbook, namely, MI (defined as notes and coins in circulation outside the banking system plus demand deposits with commercial banks). As concerning the choice f appropriate scale variables, we choose gross domestic product for income scale while choosing total consumer expenditure as consumption scale.

Regarding the price series, the ideal prices deflator would be an expenditure deflator in which the weights reflected the components of expenditure for which money is used. This deflator is however not available in Nigeria. Whereas most models of the demand for money in developed economies use the GDP or GNP deflator, we choose to use the consumer price index. In an open economy such as Nigeria, GDP deflator is not appropriate since it is constructed as a value-added deflator, which includes but excludes imports. The CPI deflator avoids this problem since it includes imports and excludes exports. And since the majority of total expenditure is on consumption, the CPI provides a reasonable first-order approximation to the true price deflator. Concerning interest rate, we note that there is only a small number of interest bearing assets held by the private sector. Throughout most of the period under examination, domestic interest rates have been controlled by the authorities. However, all interest rates have generally been adjusted in a consistent manner over the period such that despite the absence of an active market mechanism through which interest rate changes are transmitted, all the main rates of interest have tended to move together. Consequently, using discount rate (which refers to the rate at which the monetary authorities lend or discount eligible paper for depot money banks) is a reasonable approximation to the true interest rate. More, the choice of discount rate is only determined by the fact that it is the only consisted annual interest rate series. All the collected data series are reported in the appendix.

Next, we investigate the time series characteristics of our data so as to ensure consistency in subsequent econometric modeling. Table 2.1 presents evidence on the presence of unit roots in our variables using the two commonly used tests: Dickey-Fuller tests and Augmented Dickey-fullers tests which uses the regression

$$\Delta X_t = \beta X_{t-1} + U_t \qquad U_t \sim IN(0, \sigma^2) \qquad (2.10)$$

to test the null hypothesis of non – stationary for the series X_t using the t- statistic on the β parameter. The t-statistic is compared with specific values constructed by Dickey and Fuller (1979,1981) and Engle and Granger (1987) using numerical simulation methods. However, the problem is that the residuals from (2.10) should be found to be white noise. Otherwise, the equation (6.10) has to be modified to take into account higher autoregressive processes namely

$$\Delta X_t = \beta X_{t-1} + \sum^{n}_{i=1} \Delta X_{t-1} + U_t \qquad (2.11)$$

Where the n is chosen large enough so as to ensure that the residuals are white noise. The t-statistic on β in (2.11) is used to implement an augmented Dickey Fuller test (ADF) which is also reported in table 2.1 for the variables shown. Looking at thee levels of the variables, there is (not surprisingly) strong evidence in favour of the null hypothesis of non – stationarity. All the test statistics (absolute values) were lesser than the critical values at 5% and 10% significant levels. But turning to the differences of the variables, the tests overall provides support to reject the null hypothesis of non-stationarity of the series, Leading us to conclude that all the original series seem to be I(1)

TABLE 2.1: UNIT ROOTS TESTS

VARIABLE X	**UNIT ROOT IN X**		**VARIABLE ΔX**	**UNIT ROOT IN ΔX**	
	DF	*LAG LENGTH*		*DF*	*LAG LENGTH*
LM1	-1.5866	0	**ΔLMI**	-3.9054*	0
LINT	-2.1035	0	**ΔLINT**	-6.8968*	0
LP	0.84696	0	**ΔLP**	-3.7000*	0
RLGDP	-1.1954	0	**ΔRLGDP**	-4.6687*	0
RLC	-1.4130	0	**Δ RLC**	-4.2482*	0
	ADF	LAG LENGTH		ADF	LAG LENGTH
LM1	-2.2418	2	**Δ LMI**	-3.3805**	1
LINT	-1.7152	2	**Δ LINT**	-4.4521*	1
LP	1.1948	2	**Δ LP**	-5.0471*	1
RLGDP	-1.3614	2	**Δ RLGDP**	-3.7047*	1
RLC	-1.1776	2	**Δ RLC**	-5.1422*	1

** indicates statistical significance at **5%** level*
*** indicates statistical significance at **10%** level*
95%** critical value for the Augmented Dickey-fuller statistic = **-3.55
90%** critical value for the Augmented Dickey-fuller statistical = **-3.18

3.0 COINTEGRATION AND DYNAMIC ANALYSIS

Given that statistic underpinning of modern time series analysis require data to be covariance stationary, and that most macroeconomic series display significant trends has led to first difference time series before estimating economic models. Such an approach, however removes much of the long-run characteristics of the data. Engle and Granger (1987) noted that even though economic series may wander through time, economic theory often provides a rationale why certain variables should obey equilibrium constraints. That is, there may exist some linear combination of the variables that, overtime, converges to equilibrium. If the separate economic series stationary only after differencing but a linear combination of their levels is stationary, then the series are said to be cointegrated. However, this test for cointegration proposed by Engle and Granger does not distinguish between the existences of one or more cointegrating vectors. More importantly, their test relies on a super convergence result and applies an OSL estimates to obtain estimates of the cointegrating vector. These OLS estimate in practice will differ with the arbitrary normalization implicit in the selection of the left-hand-side variables for the regression equation; and moreover, different arbitrary normalization's can in practice alter the Engle and Granger test results.

In contrast, Johansen (1988), and Johansen and Juselius (1990) provide a procedure to examine the question of cointegration in a multivariate setting. However, stock and Watson (1988) provide alternative estimate based on a principal components estimation approach. Johansen approach yield maximum likelihood estimators, of the unconstrained cointegrating vector, and allows one to explicitly test for number of cointegrating vectors. This approach does not rely on an arbitrary normalisation; and test of certain restrictions suggested by economic theory, such as sign and size of estimated elasticities may also be conducted (see Hafer and Jansen, 1991; Nwaobi, 1993a; Nwaobi, 19993b). Following this approach, consider

$$X_t = \prod_1 X_{t-1} + \ldots + \prod_k X_{t-K} + \in_t \quad (t=1,\ldots T) \tag{3.1}$$

Where X_t is a sequence of random vectors with components $(X_{1t},\ldots X_{pt})$. The innovations in this process, $\in_1,\ldots,\in_T$, are drawn from a p-dimensional i.i.d Gaussian distribution with covariance Δ, and $X_{-k+1},\ldots,X_0$ are fixed. Because most economic variables are non-stationary in their levels, vector autoregressive models such as (3.1) generally are estimated in first-difference form. While such an approach satisfies the requirement that the data are stationary, it also implies some loss of information if the series are cointegrating. Letting Δ represent the first difference operator, (3.1) could be written in the equation form

$$\Delta X_t = \Gamma_1 \Delta X_{t-1} + \ldots + \Gamma_{k-1} \Delta X_{t-K+1} \Pi X_{t-k} + \in_t \tag{3.2}$$

where $\quad \Gamma 1 = -1 + II_1 + \ldots + \Pi i \; (i=1,\ldots,k-1)$ and

$$\Pi = 1 - H_1 - \ldots - \Pi_k \tag{3.3}$$

Equation (3.2) is derived by first subtracting X_{t-1} from both sides of (3.1) and collecting terms on $X_{t.1}$. The zero is added to the RHS of the equation; that is, add $-(\Pi_t-1)X_{t-1}+(\Pi_1-1)\ X_{t-1}$. Repetition of this procedure and collection of terms yields equation (3.3). The only difference between a standard first difference version of a vector auto-regressive model and (3.2) is the term ΠX_{t-k}. It is this Π matrix that conveys information about the long-run relationship between the X variables. If X_t is non-stationary in levels but ΔX_t is stationary, then X_t is integrated of order one. The individual elements of X_t may be cointegrated, however, so that one or more linear combinations of these non-stationary elements are stationary.

Cointegration can be detected by examining the $\prod$ matrix. If P X P matrix $\prod$ has rank O then all elements of X_t have roots and first differencing might be recommended. If $\prod$ is of full rank p, then all elements of X_t are stationary in levels. If the rank of $\prod$ denoted, as r is o, then there are p stochastic trends among the p elements of X. That is, all elements are non-stationary and so are all linear combinations of these elements. Likewise, if r = p, then there are p linear combinations of the p elements of X that are stationary. However, these p linear combinations span the entire space of X, so each elements of X is stationary. The interesting case in this study is when $0 < \text{rank}(\prod) = r < p$. In this case, it is said that there are cointegrating relations among the elements of X_t, and p-r common stochastic trends. If $\prod$ has rank $r < p$, this implies that $\prod = \alpha\beta^{l}$, where α and β is interpreted as a matrix of cointegrating vectors, and α is a matrix of error correction parameters. Now consider multiplying (6.13) by the matrix β where $\beta^{l} = (\beta\alpha_1)$ and α_1 is a p x (p-r) matrix orthogonal to α. This multiplication yields

$$\{\beta' \alpha'_1\} \Delta X_t = \{\beta' X'_1\} [\Gamma_1 \Delta X_{t}-_1 + \ldots + \Gamma_k - 1 \Delta_{t-k+1} + \in_t] - \{\beta' \alpha \beta' X_{t-k} 0\} \tag{3.4}$$

Where the parameters in { } are column vectors. Thus, βX_t is a column of r stationary process-the cointegrating linear combinations of the elements of X_t (p-r) common stochastic trends of non-stationary process.

Johansen and Juselius (1990) demonstrate that β, the cointegrating vector can be estimated as the eigen vector associated with the r largest statistically significant eigen values found by solving

$$|\lambda S_{kk} - S_{ko} S_{00}^{-1} S_{ok}| = 0 \qquad (3.5)$$

where S_{00} = the residual moment matrix from a least squares regression on ΔX_t on $\Delta X_{t-1}, \ldots, \Delta X_{t-k+1}$

S_{kk} = the residual moment matrix from a least squares regression of ΔX_{t-k} on ΔX_{t-k+1}

S_{ok} = is the case-product moment matrix

Using these eigen values, one may test the hypothesis that there are at most r cointegrating vectors by calculating the likelihood test statistic

$$(-2) \text{ In } (Q) = -T \, \Sigma^{P}_{r+1} \text{ In } (1-{}^{\wedge}\lambda_i) \qquad (3.6)$$

where ${}^{\wedge}\lambda_{r+1}, \ldots, {}^{\wedge}\lambda_p$ are the p-r smallest eigen values. This test was called trace test. They also develop a likelihood ratio test called the maximal eigen value test. In that test the nulls hypothesis of cointegrating vectors is tested against the alternative of r+1 cointegrating vectors. Tables (3.1) and (3..2) reports the cointegrating test results of this paper.

TABLE 3.1

TESTING FOR THE NUMBER OF COINTEGRATING VECTORS(r) ASSUMING NO INTERCEPTS OR TRENDS

(A) TEST BASED ON MAXIMAL EIGENVALUE AND TRACE OF THE STOCHASTIC MATRIX

H_O: Null Hypothesis	**H_1:** Alternative hypothesis	**Maximal** Eigen Values	**95%** Critical Value	**90%** Critical value	**Trace** Statistics	**95%** Critical Value	**90%** Critical value
r =0	r=1	17.6751	23.9200	21.5800	43.6879	39.8100	36.6900
r<=1	r=2	15.7689	17.6800	15.5700	26.0127	24.0500	21.4600
r<=2	r=3	9.0708	11.0300	9.2800	10.2438	12.3600	10.2500
r<=3	r=4	1.1730	4.1600	3.0400	1.1730	4.1600	3.0400

(B) ***TEST USING MODEL SELECTION CRITERIA***

RANK	LL	AIC	SBC	HQC	
r=0	80.6378	64.6378	52.4269	60.4735	
r=1	89.4754	66.4754	48.9222	60.4892	
r=2	97.3598	69.3598	47.9908	62.0724	
r=3	101.8952	70.8952	47.2366	62.8270	
r=4	102.4817	70.4817	46.0600	62.1532	

LL⇒*MAXIMIZED LOG-LIKELIHOOD*
AIC⇒*AKAIKE INFORMATION CRITERION*
SBC⇒*SCHWARZ BAYESIAN CRITERION*
HQC⇒*HANNAN-QUINN CRITERION*

TABLE 3.2

TESTING FOR THE NUMBER OF COINTEGRATING VECTOR(r) ASSUMING UNRESTRICTED INTERCEPTS AND NO TRENDS

(A) TEST BASED ON MAXIMAL EIGENVALUE AND TRACE OF THE STOCHASTIC MATRIX

H_O: Null Hypothesis	**H_1:** Alternative hypothesis	**Maximal** Eigen Values	**95%** Critical Value	**90%** Critical value	**Trace** Statistics	**95%** Critical Value	**90%** Critical value
r =0	r=1	17.8184	27.4200	24.9900	42.6959	48.8800	45.7000
r<=1	r=2	12.5090	21.1200	19.0200	24.8775	31.5400	28.7800
r<=2	r=3	9.9267	14.8800	12.9800	12.3685	17.8600	15.7500
r<=3	r=4	2.4418	8.0700	6.5000	2.4418	8.0700	6.5000

(B) TEST USING MODEL SELECTION CRITERIA

RANK	LL	AIC	SBC	HQC	
r=0	85.0816	65.0816	49.8180	59.8763	
r=1	93.9908	66.9908	46.3849	59.9636	
r=2	100.2453	68.2453	43.8236	59.9168	
r=3	105.2087s	70.2087	43.4974	61.0993	
r=4	106.4296	70.4296	42.9551	61.0600	

LL⇒*MAXIMIZED LOG-LIKELIHOOD*
AIC⇒*AKAIKE INFORMATION CRITERION*
SBC⇒*SCHWARZ BAYESIAN CRITERION*

HQC⇒HANNAN-QUINN CRITERION

Irrespective of which set of critical values one uses, there is a clear conflict between the test results based on the maximum eigen value statistic and the trace statistic. Assuming no intercepts or trends in the model, the maximum eigen value statistic does not reject r = 0, while the trace statistic does not reject r = 2 at the 95% percent significant level. Changing the significant level of the two tests to 90 percent results in the maximum eigen value statistic selecting r = 2, and in the trace statistic selecting r = 3. Turning to the model selection criteria, we find that the Akaike Information Criterion (AIC) and Schwarz Bayesian Criterion (SBC) choose r = 4 and r = 0 respectively, while Hannan – Quinn Criterion (HQC) choose r = 4. Alternatively assuming unrestricted intercepts and no trends in the model, the maximum static does not also reject r = 0 while trace statistics does not also reject r=0 at the 95 percent significant level. And changing the significance level of the two tests to 90 percent equally yields the same results. However,, using the model selection criteria, we again find that the AIC and SBC choose r = 4 and r = 0 respectively, while HQC chooses r=4.

Our data in this application therefore seems inconclusive on the appropriate choice of r. But turning to the long-run economic theory, we would expect two cointegrating relations. This is because cointegrtating vectors can be thought of representing constraints that an economic system imposes on the movement of the variable in the system in the long-run. Consequently, the more cointegrating vectors there are, the "more" the system. The fewer the number of cointegrtating vectors, the less constrained is the long run relationship. Hence, it is desirable to have many cointegrating vectors, since it ensures an economic system to be stationary in as many directions as possible. That is, we prefer economic models that have unique steady-state equilibria,. Table 3.3 reports the Π (estimated long-run) matrix corresponding to equation (3.2) while the eigen vectors (β) are presented in table 3.4.

TABLE 3.3 ESTIMATES OF Π

	LM1	**RLGDP**	**LP**	**LINT**
LM1	-0.43494	0.69029	0.44438	0.10920
RLGDP	-0.23044	0.24066	0.26052	-0.052218
LP	0.16575	-0.042839	-0.021350	0.15220
LINT	0.37175	-0.38428	-0.42106	0.087713

TABLE 3.4: β^{l} MATRIX (COINTEGRATED VECTOR)

	LMI	***RLGDP***	***LP***	***LINT***
ECMI	-0.97974	0.96802	1.1187	-0.27056
	(-1.0000)	(0.98803)	(1.1418)	(-0.27615)
ECM2	0.51439	-1.1095	-0.46679	-0.38711
	(-1.0000)	(2.1569)	(0.90747)	(0.75256)

ECMI REPRESENTS THE FIRST COINTEGRATING VECTOR
ECM2 REPRESENTS THE SECOND COINTEGRATING VECTOR
COEFFICIENTS IN PARENTHESIS ARE NORMALISED.

In order to impose the cointegrating vectors on the error correction model, we regress

$$A(L)\,\Delta LM1_t = \delta o + \Delta RLGDP_t + C(L)\Delta LP_t + D(L)\Delta LINT_t + E(L)ECM1_{t-1} + F(L)ECM_{t-1} + \sum \alpha_i D_i + \in_t \qquad (3.7).$$

Where A(L) … F(L) are polynomials of the form $A(L) = \sum \phi_i L^i$ in which L is the lag operator such that $L^r X_t = X_{t-r}$ and D_i are dummy variables. ECM is the Error Correction Vector. Equation (3.7) can be written in a more general form as

$$\Delta LM1_t = \delta o + \phi_1 \Delta LM1_{t-1} + \phi_2 \Delta LM1_{t-2} + \beta_1 \Delta RLGDP_t + \beta_2 \Delta RLGDP_{t-1} + \beta_3 \Delta RLGDP_{t-2} + \lambda_1 \Delta LP_t + \lambda_2 \Delta LP_{t-1} + \lambda_3 \Delta LP_{t-2} + \theta_1 LINT_t + \theta_2 \Delta LINT_{t-1} + \theta_3 \Delta LINT_{t-2} + \prod_1 ECM1_{t-1} + \prod_2 ECM2_{t-1} + \alpha_1 SAD + \alpha_2 WAD + e_t \qquad (3.8)$$

The results of the regression equation (3.8) are presented as equation (3.9):

$$\Delta LM1_t = \qquad (3.9)$$

	Δ**LM1$_{t-i}$**	Δ**RLGDP$_{t-I}$**	Δ**LP$_{t-I}$**	Δ**LINT$_{t-i}$**
i = 0		0.55248	0.32351	-0.19604
		(2.9524)	(1.0459)	(-1.3466)
i = 1	0.12424	-0.12179	0.11563	-0.13410
	(0.47118)	(-0.34981)	(0.33695)	(-0.61421)
i = 2	0.093202	-0.53948	-0.72482	-0.11294
	(0.45266)	(-1.7536)	(-1.6387)	(-0.49289)

CONSTANT	**ECM1$_{t-1}$**	**ECM2$_{t-1}$**	**SAD**	**WAD**
-3.9448	0.15387	-0.62574	-0.055283	0.0028903
(-4.0482)	(0.68375)	-3.5552	(-0.67557)	(0.26558)

[Values in parenthesis are estimated t ratios; T=1960 – 1995, R^2 = 0.79529, = σ0.11582, F(15,17) = 4.4030, DW = 2.5584;
ξ_I(I) = 10.2303, ξ_1 (1,16) = 7.1887, ξ_2 (I) = 0.30172, ξ_2 (1,16) = 0.14764, ξ_3(2) = 0.71147,
ξ_4 (1) = 0.00350, ξ_4 (1,31) = 0.00329].

$\xi_1 \Rightarrow$ Lagrange multiplier test of residual serial correlation (x^2 and F versions).
$\xi_2 \Rightarrow$ Ramsey's reset test using the square of the fitted values.
$\xi_3 \Rightarrow$Normality test based on a test of skewness and Kurtosis of residuals
$\xi_4 \Rightarrow$ Heteroscedasticity test based on the regression of squared residuals on fitted values.

The regression results (3.9) satisfies the diagnostic checks reported and provides generally sensible estimates, despite very high intercorrelations between regressors as well as the presence of serial correlation. Given the over-parameterization of equation (3.8), a sequential simplification search (similar to general to specific approach) is then undertaken to reach a more parsimonious model. The representation in (3.9) was therefore simplified to the error-correction model (3.10) by transforming the model to an interpretable and near orthogonal specification and eliminating negligible and insignificant effects:

$$\Delta LMI_t = \underset{(-3.6816)}{-2.1343} + \underset{(4.3309)}{0.63910}\Delta RLGDP_t + \underset{(2.2824)}{0.41295}\Delta LP_t - \underset{(0.88935)}{0.097526}\Delta LINT_t - \underset{(-3.7740)}{0.36165}ECM2_{t-1} \qquad (3.10)$$

[Values in parenthesis are estimated T-ratios; T = 1960 – 1995, R^2 = 0.63, $\hat{\sigma}$= 0.12002, F(4,30) = 12.8721, DW = 1.9020; $\xi_1(1)$ = 0.056984, $\xi_1(1,29)$ = 0.047292, $\xi_2(1)$ = 0.70539, $\xi_2(1,29)$ = 0.59649, $\xi_3(2)$ = 0.185611, $\xi_4(1)$ = 1.6123, $\xi_4(1,33)$ = 1.5935, $\eta 1(4)$ = 1.9640, $\eta 1(4,26)$ = 0.38642 $\eta 4(4)$ = 5.5577 $\eta 4(4,26)$ = 1.2270]

ξ^1s $\Rightarrow$ As already defined above.

H_1 $\Rightarrow$ Higher order test of serial correlation of residuals

η_4 $\Rightarrow$ Autoregressive conditional heteroscedasticity of residuals (Arch)

Equation (3.10) is similar in form and in numerical parameter values to several successful money demand models for the developing countries (see Arestis and Demetriades, 1991; Adams, 1992; Nwaobi, 1993a: and Choudhry, 1995). Its coefficients satisfy the sign restrictions on the equation to be interpretable as a money-demand function. The coefficients sizes imply large immediate responses to changes in income and prices but slow adjustment subsequently to interest rates and remaining disequilibria, via the error correction term. The empirical parameterization in (3.10) exhibits multiple equilibria, with two corresponding to the long-run solution (3.9). Concerning the statistical attributes of (3.10), the various diagnostic checks are insignificant (if regards as test statistics) and indicate design of a model congruent with the information available. From the reported diagnostic tests, the residuals are white noise. There is no ARCH, RESET, or heteroscedastic evidence of misspecification; the residuals are approximately normally distributed.

Any claim to the constancy of the model would however need both constant parameters and adequacy of predictions. The predictive failure test (Chows second test), which is a test of adequacy of predictions, was therefore carried out. For this purpose, our model was re-estimated over the sample period, 1960 – 1990, and the resultant required predictive failure test statistics (ξ_5) was given as ξ_5 (1) = 2.8235 and ξ_5 (5,25) = 0.5647. We therefore accept the null hypothesis of predictive pass, which suggests that our model can forecast with minimum errors as shown in table 3.5.

TABLE 3.5 FORECAST PERFORMANCE

OBSERVATION	ACTUAL VALUES	FORECAST VALUES	ERRORS
1991	0.34373	0.40423	-0.060504
1992	0.44239	0.50060	-0.058214
1993	0.42774	0.27724	0.15049
1994	0.38747	0.51268	-0.12522
1995	0.15651	0.34116	-0.18466

MPE (MEAN PREDICTION ERRORS) = -0.055619
SSM (SUM SQUARES PREDICTION ERROR) = 0.015895
MSE (MEAN SUM ABSOLUTE PREDICTION ERRORS) = 0.11582
RSM (ROOT MEAN SUM SQUARE PREDICTION ERRORS) = 0.12608
PREDICTIVE FAILURE TEST F(5,25) = 0.56471.

Furthermore, Lucas (1976) stresses the instability of the parameters under different policy regimes and structural changes. Thus, the stability of the parameters of our model are examined using the plot of cumulative sum (cusum) and the cumulative sum of squares (cusumsq) of residuals. The cusum test is used primarily with recursive residuals, because OLS residuals suffer from the constraint that the residuals finally sum to zero. The cusum is, in fact, the estimated standard error of the recursive residuals times the summation of them. Thus, if there is any structural changes or mis-specification of the model, the residuals will show up to have same signs. The cumulative sum (cusum) test is described in Brown et. al. (1975) and is based on the cusum of recursive residuals defined by

$$Wr = 1/\hat{\sigma}_{ols}\Sigma^{r}_{j=k+1} (V_j), \quad r = k + 1, k + 2, \ldots\ldots, n \qquad (3.11)$$

where V_t is the recursive residual based on the first j observations given by $V_r = (y_r - X'_r\, \beta_{r-1})/dr$, $r = 0\ k + 1, k + 2, \ldots n$ and where β_r are defined by $\beta_r = (X_r{}'X_r)^{-1}X'_r y_r$, $r = k + 1, k + 2, \ldots, n$ and $dr = \sqrt{\{1 + X'_r (X'_{r-1}X_{r-1})^{-1}X_r\}}$ and also where $\hat{\sigma}_{ols}$ is defined as the standard error of the regression given by $\hat{\sigma}^2_{ols} = (y - X\hat{\beta}_{ols})'(y - X\hat{\beta}_{ols})\ (n - k)$
This test employs a graphic technique and involves plotting Wr and a pair of straight lines for values of $r = k + 1, k + 2, \ldots n$. The straight lines are drawn assuming a five percent significance level. The equations of the line are given by $W = \pm[0.948 \sqrt{(n-k)} + 1.896(r-k)\sqrt{(n-k)}]$ for $r = k + 1, k + 2, \ldots, n$.
On the other hand, the CUSUM of squares test employs the squared recursive residuals, V_j^2. It is based on the quantities

$$WW_r = \Sigma^{r}_{j=k+1} (V_j^2) / \Sigma^{n}_{j=k+1} (V_j^2), \quad r = k + 1, k + 2, \ldots\ldots, n \qquad (3.12)$$

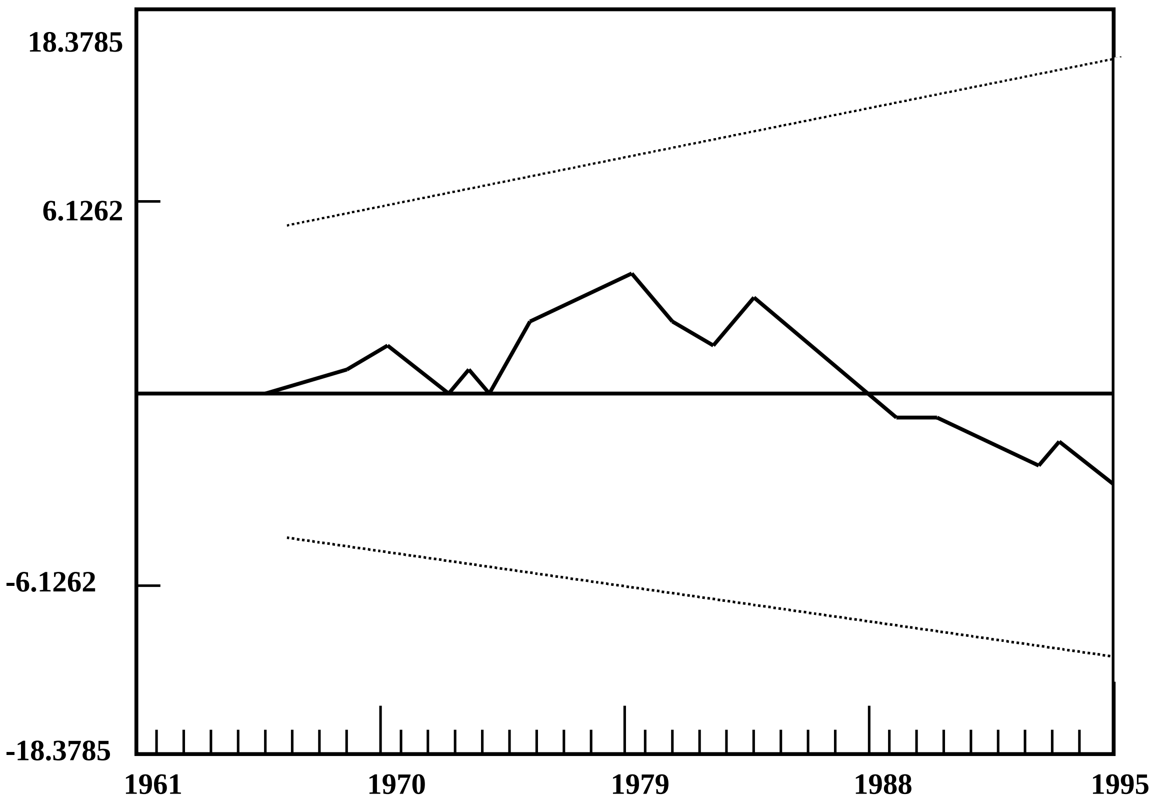
and involves plotting WWr and a pair of lines whose equations are given by

$$WW = \pm\ C_o + (r - k) / (n - k), \quad r = k + 1, k + 2 \ldots\ldots, n \qquad (3.13)$$

Where C_o is determined by the significance level chosen for the test. In other words, the cusumq test is based on a plot of the ratio of the squares of the residuals of the test period against the squares of the residuals of the whole sample. Two lines are drawn above and below to provide a means of stressing the significance of departures. We reject the null hypothesis of no structural change, if either of the two lines is crossed and vice versa. Figures 3.1 (cusum plot) and figure 3.2 (cusumsq plot) show the stability tests for the estimated dynamic money demand equation (3.10).

FIGURE 3.1

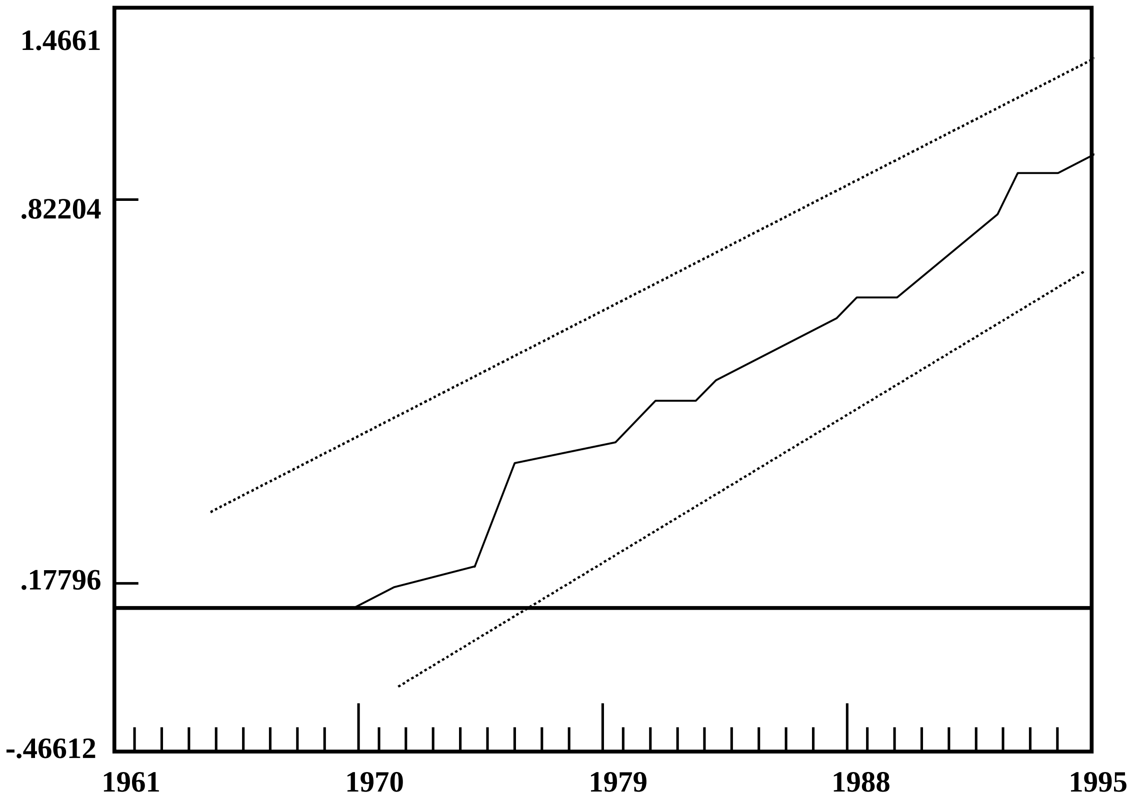
Plot of Cumulative Sum of Recursive Residuals



The straight lines represent critical bonds at 5% significance level

FIGURE 3.2

Plot of Cumulative Sum of Squares of Recursive Residuals



The straight lines represent critical bonds at 5% significance level

From these graphical plots, we conclude that our equation is stable throughout the period under analysis and this result is consistent with the ARCH test already presented above.

At this juncture, and inline with the objectives of the study, we re-estimated our dynamic money demand equation using consumption, as the appropriate scale variable. Indeed, we adopted the same procedure as that used in estimating the final error correction model (vector) of the income-scaled money demand function. And for space limitation, we report only the estimated final equation as shown below:

$$\Delta LM1t = -1.6772 + 0.79790\Delta RLC + 0.96882\Delta LP - 0.10383\Delta LINT - 0.29282ECM11_{t-1} \quad (3.12)$$
$$\quad\quad (-2.2912) \quad (4.2444) \quad\quad (5.3710) \quad\quad (-0.58900) \quad\quad (-2.3780)$$

[Values in parenthesis are estimated t-ratios; R^2 = 0.558, T = 1960 – 1995, σ = 0.13250, F(4,28) = 8.8626, DW = 2.1486; ξ1(1) = 0.3892), ξ1(1,27) = 0.32267, ξ2(1) = 0.98949, ξ2(1,27) = 0.83461, ξ3(2) = 0.0065911, ξ4(1) = 0.10076, ξ4(1,31) = 0.04945]

Again, the reported diagnostic tests supports a sensible consumption scale-money demand function of the Nigerian Economy. But comparing it to income scale-money demand function will result in the selection of the income as the more appropriate scale for the Nigerian case.

4.0 NONNESTED MODELLING (ENCOMPASSING)

The theory of encompassing offers an improved empirical research strategy, specifically developed to augment the tradition of empirically testing with the further requirement that a model should be able to explain or account for the results obtained by rival models. Consider the following two linear regression models:

$$M_1 : y = X\beta_1 + u_1, \qquad u_1 \sim \quad N(0, \sigma^2 I_n) \tag{4.1}$$

$$M_2 : y = Z\beta_2 + u_2, \qquad u_2 \sim \quad N(0, w^2 In) \tag{4.2}$$

Where y is the n x 1 vector of observations on the dependent variable; X and Z are n x k_1 and n x k_2 observation matrices for the regressors of models M_1 and M_2; β_1 and β_2 are the k_1 x 1 and k_2 x 1 unknown regression coefficient vectors; and u_1 and u_2 are the n x 1 disturbance vectors. Models M_1 and M_2 are said to be non-nested if the regressors of M_1 (respectively M_2) cannot be expressed as an exact linear combinations of the regressors of M_2 (respectively M_1). For the purpose of our paper therefore, the consumption scale money demand model is labeled model 1 while the income scale-money demand model is labeled model 2.

The various kinds of non-nested tests, information and likelihood criterion are portrayed in tables 4.1 and 4.2. The N-test is attributed to Cox (1961, 1962) as modified by Persaran (1974). The NT-test is the adjusted Cox-test derived in Godfrey and Pesaran (1983). The W-test is the Wald-type test of M_1 against M_2 proposed in Godfrey and Pesaran (1983) The J-test is due to Davidson and Mackinnon (1981). This test is valid asymptotically, but in small samples the NT-test, and W-test are preferable to it. The JA-test is due to Fisher and Mcaleer (1981). The Encompassing test was proposed by Deaton (1982), Dastoor (1983), and Mozon and Richard (1986). In the case of testing M_1 against M_2, the encompassing test is the same as the classical F-test and is computed as the F-statistic for testing $\delta = 0$ in the combined OLS regression.

$$y = X_{ao} + Z^*\delta + u \tag{4.3}$$

where Z* denotes the variables in M_2 that cannot be expressed as exact linear combinations of the regressors of M_1. This is encompassing test is asymptotically equivalent to the above non-nested tests under the null hypothesis, but in general it is less powerful than these for a large class of alternative non-nested models. The choice criteria or model selection criteria are Akaika Information Criterion (Akaike, 1973) and Schwartz Bayesian Information 'Criterion (Schwartz, 1978). Both use statistics, which incorporate measures of the precision and parsimony in parameterization of models.

Four other non-nested test statistics and two choice criteria are used for pair-wise testing and choice between linear, log-linear and ratio models. The PE-test statistic is proposed by Mackinnon et. Al (1983). The BM-test statistic is proposed by Bera and McAleer (1989) and is used for testing linear versus log-linear models. The Double-Length (DL) regression statistics is proposed by, Davidson and Mackinnon (1984). The Simulated Cox test statistics, denoted by S-test is developed by Pesaran and Pesaran (1993) and subsequently applied to tests of linear versus log-linear models, and first-difference versus log-difference stationary models (Pesaran and Pesaran, 1995). Sargan's (1964) likelihood criterion compares the maximized values of the log-likelihood functions under M_1 and M_2; while Vuong's criterion (1989) is motivated in the context of testing the hypothesis that M_1 and M_2 are equivalent, using the Kullback Leibler information criterion as a measure of goodness of fit.

TABLE 4.1 TESTS FOR NON-NESTED REGRESSION MODELS.

	HO: MODEL 1	**HO: MODEL 2**
	H1: MODEL 2	**H1: MODEL 1**
N-TEST	-4.7688	-3.2303
NT-TEST	-3.1782	-2.1028
W-TEST	-2.6258	-1.8589
J-TEST	3.3964	2.5626
JA-TEST	1.8419	1.4605
ENCOMPASSING F(3,25)	3.7696	2.2986 F(3,25)

AIC (MODEL 1 VERSUS MODEL 2) = -2.1381 FAVOUR MODEL 2
SBIC (MODEL 1 VERSUS MODEL 2) = -2.1381 FAVOUR MODEL 2

TABLE 4.2 NON-NESTED TESTS BY SIMULATION.

	M1 AGAINST M2	**M2 AGAINST M1**
S-TEST	-2.4929	-1.5062
PE-TEST	3.3964	2.5626
BM-TEST	1.8419	1.4605
DL-TEST	3.4330	2.7304

SLC (MODEL 1 VERSUS MODEL 2) = -2.1381 FAVOURS M2
VLC (MODEL 1 VERSUS MODEL 2) = -4.4447 FAVOURS M2

AIC = AKAIKE INFORMATION CRITERION
SBIC = SCHWARTZ BAYESIAN INFORMATION CRITERION
SLC = SARGEN'S LIKELIHOOD CRITERION
VLC = VUONG'S LIKELIHOOD CRITERION

Looking at table 4.1 (the non-nested regression models tests), we observe that the evidence adduced therein is overwhelmingly in favour of model 2, in which income is the scale variable. All the six test statistics clearly reject model 1 against the alternative of model 2. at the same time, in no case did the null hypothesis that model 2 is true be rejected model 1. Also, the two information criteria (AIC and SBIC) reported clearly favours model 2. All the four tests in the various replications clearly reject model 1 in favour of model 2. Equally, the two-likelihood criterion (SLC and VLC) reported clearly favours model 2 as the acceptable demand for money function in Nigeria. These evidence provides strong support to the proposition that income is a better scale variable than consumption scale variable in modeling the demand for money function in Nigeria.

5.0 SUMMARY AND CONCLUSIONS

This paper has addressed the estimation and testing problem of long-run relations in economic modeling. With a relatively simple model-specifying vector valued autoregressive process, the hypothesis of the existence of cointegration vectors is formulated as the hypothesis of reduced rank of the long-run impact matrix. This is given a simple parameter form, which allows the application of the method of maximum likelihood and likelihood ratio tests. In this way, we derived estimates and test statistics for the hypothesis of a given number of cointegration vectors as well as estimates and tests for linear hypothesis about the cointegration vectors.

Using Nigerian data, we found that the demand for money (LM1) is cointegrated with real income (RLGDP), interest rate (LINT) and price level (LP). The four variables were found to be integrated of order one. That is, they are I(1). This implied that the levels of these variables are differenced once to achieve stationarity, and applying the Johansen maximum likelihood estimation procedure, we accepted the alternative hypothesis of two cointegrating vectors. Adopting general to specific approach, an over parameterized dynamic money demand function was estimated. By eliminating negligible and insignificant effects, the final congruent and parsimonious money demand equation was arrived at. This money demand function was found to be stable and the evidence was seen from the plots of cusum and cusumsq tests. Thus, that there appears to be stable long-run relationship between real money, real income and nominal interest rates establishes the potential for achieving price level stability by controlling the growth of money balances.

Furthermore, evidence gathered from the non-nested tests, suggest that income is the more appropriate scale variable in the estimation of the demand for money in Nigeria. This results sharply contradict most findings based on developed countries studies but the results are in tune with the majority of studies that used income as the appropriate scale variable in demand for money functions estimated through techniques of cointegration and error correction mechanism. However, it is pertinent to note that cointegrating test results may be sensitive to the sample period used and we believe that cointegration is a long-run property and thus we often need long spans of data to properly test it. This leads to the conclusion that although cointegration is an exciting and potentially immensely important new tool, we must be careful in its use and application for policy making. As indication for further research, we will adopt impulse response functions as measures of the time profile of the effect of shocks on the future states of the dynamical systems (see Nwaobi, 2001).

APPENDIX (to Chapter Six)

DATA SOURCES, COLLECTION AND DEFINITIONS

{1} ***M1*** *IS DEFINED AS NARROW MONEY [MILLIONS OF NAIRA] AND IS DERIVED FROM INTERNATIONAL FINANCIAL STATISTICS YEAR BOOK (IFS LINE 34). IT COMPRISES TRANSFERABLE DEPOSITS AND CURRENCY OUTSIDE BANKS.*

{2} ***M2*** *IS DEFINED AS M1 PLUS QUASI-MONEY [MILLIONS OF NAIRA] AND DERIVED FROM INTERNATIONAL FINANCIAL STATISTICS YEAR BOOK (IFS LINE 34 + 35). QUASI MONEY REFERS TO LIABILITIES OF BANKING INSTITUTIONS, WHICH COMPRISE TIME, SAVINGS AND FOREIGN CURRENCY DEPOSITS.*

{3} ***INT*** *IS DEFINED AS DISCOUNT RATE [PER CENT PER ANNUM]. IT IS DERIVED FROM INTERNATIONAL FINANCIAL STATISTICS YEAR BOOK (IFS LINE 60).*

{4} ***CP*** *IS DEFINED AS CONSUMER PRICE INDEX [1990 = 100]. IT IS DERIVED FROM INTERNATIONAL FINANCIAL STATISTICS YEAR BOOK [IFS LINE 60]. HOWEVER, THE INDEX CONVERSION PROCEDURE WAS USED IN CONVERTING THE SERIES FROM THE YEARS 1960 – 1966 (THAT IS FROM 1985 BASE YEAR TO 1990 BASE YEAR).*

{5} ***GDP*** *IS DEFINED AS GROSS DOMESTIC PRODUCT [MILLIONS OF NAIRA]. IT IS DERIVED FROM INTERNATIONAL FINANCIAL STATISTICS YEAR BOOK (IFS LINE 99B).*

{6} ***CON*** *IS DEFINED AS TOTAL CONSUMPTION [MILLIONS OF NAIRA]. IT IS DERIVED FROM INTERNATIONAL FINANCIAL STATISTICS YEAR BOOK (IFS LINES 91F + 96F). HOWEVER, THE VALUE FOR 1995 WAS DERIVED FROM CENTRAL BANK OF NIGERIA STATISTICAL BULLETIN.*

{7} ***WAD IS*** *DEFINED AS WAR DUMMY*

{8} ***SAD*** *IS DEFINED AS STRUCTURAL ADJUSTMENT DUMMY*

{9} ***CTT*** *IS DEFINED AS THE INTERCEPT TERM.*

{10} ***TTR*** *IS DEFINED AS THE TIME TREND*

OBS	M1	M2	INT	CP	GDP	CON
1960	240.8000	295.6000	5.6000	2.4000	2400.0	2278.0
1961	243.0000	314.0000	5.5000	2.5000	2378.0	2248.0
1962	253.0000	333.0000	4.5000	2.6000	2516.0	2333.0
1963	269.0000	362.0000	4.0000	2.5000	2946.0	2623.0
1964	318.0000	431.0000	5.0000	2.5000	3145.0	2766.0
1965	328.0000	469.0000	5.0000	2.6000	3361.0	2812.0
1966	357.0000	520.0000	5.0000	2.9000	3614.0	3053.0
1967	323.0000	454.0000	5.0000	2.8000	2951.0	2567.0
1968	339.0000	522.0000	4.5000	2.8000	2878.0	2535.0
1969	447.0000	663.0000	4.5000	3.1000	3851.0	3321.0
1970	643.0000	979.0000	4.5000	3.5000	5621.0	4721.0
1971	670.0000	1042.0	4.5000	4.1000	7098.0	5721.0
1972	747.0000	1204.0	4.5000	4.2000	7703.0	6056.0
1973	788.0000	1370.0	4.5000	4.4000	11199.0	7899.0
1974	1619.0	2592.0	4.5000	5.0000	18811.0	12274.0
1975	2463.0	4035.0	3.5000	6.7000	21779.0	15926.0
1976	3728.0	5708.0	3.5000	8.3000	27572.0	18882.0
1977	5420.0	7675.0	4.0000	9.5000	32747.0	22888.0
1978	5101.0	7521.0	5.0000	11.5000	36084.0	29340.0
1979	6147.0	9849.0	5.0000	12.9000	43151.0	30810.0
1980	9227.0	14390.0	6.0000	14.2000	50849.0	36746.0
1981	9745.0	15239.0	6.0000	17.1000	50749.0	41182.0
1982	10049.0	16694.0	8.0000	18.4000	51709.0	43100.0
1983	11283.0	19034.0	8.0000	22.7000	57142.0	48946.0
1984	12204.0	21243.0	10.0000	31.7000	63608.0	54887.0
1985	13227.0	23153.0	10.0000	34.1000	72355.0	61408.0
1986	12663.0	23605.0	10.0000	36.0000	73062.0	63692.0
1987	14906.0	28895.0	12.7500	40.1000	108885.0	85724.0
1988	21446.0	38406.0	12.7500	61.9000	145243.0	122326.0
1989	26664.0	43371.0	18.5000	93.1000	224797.0	148904.0
1990	34540.0	57554.0	18.5000	100.0000	260637.0	157889.0
1991	48708.0	79068.0	15.5000	113.0000	324011.0	234960.0
1992	75810.0	125622.0	17.5000	163.4000	549808.0	424614.0
1993	116276.0	190334.0	26.0000	256.8000	701472.0	565056.0
1994	171303.0	259808.0	13.5000	403.3000	914334.0	781961.0
1995	200325.0	311580.0	13.5000	696.9000	1436649	1294441

OBS	WAD	SAD	CTT	TTR
1960	0.00	0.00	1.0000	1.0000
1961	0.00	0.00	1.0000	2.0000
1962	0.00	0.00	1.0000	3.0000
1963	0.00	0.00	1.0000	4.0000
1964	0.00	0.00	1.0000	5.0000
1965	0.00	0.00	1.0000	6.0000
1966	0.00	0.00	1.0000	7.0000
1967	1.0000	0.00	1.0000	8.0000
1968	1.0000	0.00	1.0000	9.0000
1969	1.0000	0.00	1.0000	10.0000
1970	0.00	0.00	1.0000	11.0000
1971	0.00	0.00	1.0000	12.0000
1972	0.00	0.00	1.0000	13.0000
1973	0.00	0.00	1.0000	14.0000
1974	0.00	0.00	1.0000	15.0000
1975	0.00	0.00	1.0000	16.0000
1976	0.00	0.00	1.0000	17.0000
1977	0.00	0.00	1.0000	18.0000
1978	0.00	0.00	1.0000	19.0000
1979	0.00	0.00	1.0000	20.0000
1980	0.00	0.00	1.0000	21.0000
1981	0.00	0.00	1.0000	22.0000
1982	0.00	0.00	1.0000	23.0000
1983	0.00	0.00	1.0000	24.0000
1984	0.00	0.00	1.0000	25.0000
1985	0.00	0.00	1.0000	26.0000
1986	0.00	1.0000	1.0000	27.0000
1987	0.00	1.0000	1.0000	28.0000
1988	0.00	1.0000	1.0000	29.0000
1989	0.00	1.0000	1.0000	30.0000
1990	0.00	1.0000	1.0000	31.0000
1991	0.00	1.0000	1.0000	32.0000
1992	0.00	1.0000	1.0000	33.0000
1993	0.00	1.0000	1.0000	34.0000
1994	0.00	1.0000	1.0000	35.0000
1995	0.00	1.0000	1.0000	36.0000

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