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THE DEMAND FOR MONEY IN SIERRA LEONE REVISITED

By Kelfala M. Kallon

ABSTRACT

This paper uses Johansen’s cointegration methodology to estimate Sierra Leone’s money long run demand function for the 1964-2005 period. It finds a stable long-run relationship between the quantity of real money balances and its determinants. Secondly, all the estimated coefficients have their expected signs. Additionally, as expected in economies with under-developed financial systems, Sierra Leone’s long-run money demand function is unit income-elastic and interest-rate inelastic. Thus, the study provides support for the neoclassical money demand specification. Additionally, it reaffirms the central findings of Kallon (1992).

Keywords: Sierra Leone; Money Demand, Cointegration Analysis

JEL: E41, O55, O11, N17
THE DEMAND FOR MONEY IN SIERRA LEONE REVISITED

By KELFALA M. KALLON

I. Introduction

In the standard Keynesian IS-LM model, monetary policy is more effective as a short-term macroeconomic stabilization tool when the money demand function is interest-rate inelastic. Meanwhile, fiscal policy is relatively more effective when money demand is interest-rate elastic. Monetarists claim the former to be the case while Keynesians hold the opposite view.

The debate that raged between Keynesians and monetarists over the efficacy of monetary policy as a short-term stabilization tool generated great interest in empirical money demand studies. However, data problems made it difficult to conduct such studies for most Sub-Saharan African countries. In fact, I know of only one such study for Sierra Leone (Kallon 1992), which used regression analysis to estimate the country’s money demand function for the 1966:1-1985:4 period. Since that study, there have been many advances in time-series econometrics (especially in unit-root and co-integration analysis) that might make its results suspect. For example, if the variables in that study contain unit roots, its findings could have been caused by spurious correlation (Granger and Newbold 1974). It is therefore necessary to revisit this issue using these recent econometric techniques.

Additionally, because one observes only the quantity of money supplied, not the quantity demanded, a stable monetary equilibrium is necessary in order to allow one to use the quantity of money supplied as a proxy for the quantity demanded. For this to make sense, the money demand curve must be stable (non-shifting over time). This means that any deviation of the quantity of money demanded from its long-run equilibrium level would be temporary.

However, as Nelson and Plosser (1982) have shown, most macroeconomic time series have unit roots. Thus, they can drift along without any tendency to return to long-run equilibrium. Engle and Granger (1987) have, however, demonstrated that a linear combination of two or more non-stationary variables can be stationary. That linear combination would be their cointegrating equation, which is alternatively considered as their long-run equilibrium relation. This makes cointegration analysis indispensable to empirical money demand studies.

Johansen (1988 and 1992) and Johansen and Juselius (1990) have developed a one-step method for testing for cointegration between non-station-
ary time-series, which is based on vector error-correction modeling. A key advantage of this technique, relative to Engle-Granger technique (Engle and Granger 1987), is that it allows one to test for the presence of cointegration between integrated variables, estimate their parameters, and even impose parameter restrictions in a one-step process.

This paper uses Johansen’s cointegration methodology to estimate the long-run money demand function for Sierra Leone. It is organized as follows. Section II summarizes its theoretical underpinnings while the data used in the analysis are described in Section III. The Johansen cointegration test is performed in Section IV. Diagnostic tests for model specification and performance are presented in Section V. Finally, a summary of the paper’s key findings and policy implications are provided in Section VI.

II. The Theory of Money Demand

Neoclassical money demand theory holds that the nominal quantity of money demanded depends on the price level \( p \), real income \( y \), and the opportunity cost of holding money \( c \). Assuming a money-bond dichotomy, the opportunity cost of holding any given monetary aggregate is the difference between the rate of return on bonds \( r_b \), and its own-rate of return \( r_m \). Additionally, as the price level rises, the purchasing power of money falls. The resultant inflation-induced loss of purchasing power is an additional cost of holding money.

At least two-thirds of the Sierra Leonean population lives in rural areas without access to formal financial institutions. Therefore, they generally hold their wealth in currency and real property—such as real estate and/or inventories of foodstuff (Kallon 1992). Hence, they hedge against inflation by investing in such real assets. Therefore, an increase in the price level leads to a substitution of real money balances with such real assets.

Additionally, the inaccessibility of modern financial institutions to rural Sierra Leoneans inhibits them from getting bank loans. Consequently, they are forced to finance their investments from their savings, which are held in real cash balances. As a result, real money balances serve as a conduit for capital accumulation. Hence, the quantity of real money balances demanded tends to increase with the ratio of desired investment to income (McKinnon 1973: 59). Therefore, ceteris paribus, as the price level rises, thus lowering the real interest rate, the quantity of capital stock demanded rises. Accordingly, the desired capital stock rises. Accordingly, the quantity of real money (currency) balances demanded by economic agents to satisfy this “investment” motive also rises. Thus, the net impact of inflation on the demand for money in countries with under-developed financial sectors such as Sierra Leone’s will be negative if the negative “hedging” effect outweighs the positive “investment” effect, and vice versa.
Finally, prior to the financial reforms of the mid-1980s, it was illegal for Sierra Leoneans to hold foreign currency. All such currencies were to be surrendered to the monetary authorities for balance-of-payments purposes. The reforms of the mid-1980s got rid of this restriction and allowed Sierra Leoneans to hold foreign-currency accounts—even within the country. Furthermore, the nation’s currency, the leone, was floated.

Foreign-currency-denominated assets earn the foreign interest rate \( r^* \). They also provide protection against future depreciations of the domestic currency. Thus, the expected return on foreign-currency-denominated assets is the sum of the interest rate on foreign bonds \( r^* \) and the expected rate of depreciation of the domestic currency \( \Delta x_e \). However, holding foreign-currency-denominated bonds implies foregoing the interest income from domestic bonds \( r_b \). Thus, assuming uncovered interest-parity, the decision to allocate one’s wealth between domestic financial assets and their foreign counterparts is positively related to the spread \( s \) between the returns to the former and the latter. In other words, domestic residents will increase their holdings of domestic financial assets (inclusive of bonds and real money balances) when \( r_b > r^* + \Delta x_e \), and vice versa.

Based on the assumption of unit-price homogeneity, the demand for real money balances can be stated as follows:

\[
 m^d = f(y, r_b-r_m, s, \Delta p); \quad (1)
\]

where \( m^d \) and \( \Delta p \), respectively, denote real money balances demanded and the inflation rate, while the other variables are as defined earlier.

III. The Data

The data used in this study, which span the 1964-2005 period, are from the June 2007 issue of the International Monetary Fund’s (IMF) *International Financial Statistics on CD-ROM* (IMF 2007). Money (M) is the narrow measure of money (M1) while the consumer price index (2000=100) represents the price level. Because M1 has no yield, its own-rate of return is zero. Hence, the opportunity cost of holding M1 is the rate of return on domestic bonds, which is represented by the treasury-bill rate. The US treasury-bill rate was used as the appropriate foreign interest rate. This is justified by the fact that the US dollar has been a parallel currency in Sierra Leone since the leone was floated and Sierra Leoneans started using the dollar as a hedge against exchange-rate risk.

Although real gross national income (GNI) measures income in an open economy more accurately than real gross domestic product (GDP), real GDP (in 2000 constant dollars) is used as the scale variable here because it has three more observations than real GNI.

Unfortunately, IMF (2007) does not report the GDP deflator (2000=100) for years prior to 1971. However, the *International Financial Statistics*

Because the expected rate of depreciation of the leone is unobserved, the log of the nominal exchange rate was regressed on its one-period lag, a trend term, and a third- and fourth-order moving-average error terms. The first difference of the fitted values from this regression equation was used as a proxy for the expected change in the exchange rate. This was then used to calculate the interest-rate spread between leone- and dollar-denominated bonds.

Figures 1A and 1B are graphs of real money balances (m) in levels and first-difference, respectively. The first-difference series is plotted around its mean and 1.96 standard-error bands. In levels (Figure 1A), there is a slight positive trend in the real money stock up to 1980. Thereafter the trend was reversed until 1996 when it became positive again. Moreover, as expected, there is no visible trend in its first-difference. However, there is a large outlier at 1987 where it falls outside the lower band.

Real GDP and its first difference (with 1.96 standard-error bands) are plotted in Figures 2A and 2B, respectively. There is a slight upward trend in real GDP up until 1990. During the war years, this trend was reversed. It became positive again in 2000 and has remained so since. With the exception of 1992, 1997, and 2002, the first difference of real GDP stayed within the 1.96 standard-error bands practically throughout the sample period.

The interest rate was controlled throughout the pre-reform period (1964-1985). This is reflected in Figures 3A and 3B, which are graphs of the treasury-bill rate and its first-difference. Figure 3A shows that the post-reform market correction caused the treasury-bill rate to shoot up to high double-digits between 1986 and 2000. It then fell to low double-digits. Consequently, there is relatively no variation in its first-difference during the 1964-1985 period. Additionally, it fell outside the 1.96 standard error bands only thrice during the sample period (in 1990, 1992, and 1993).
Figure 1A:
Real Money Stock

Figure 1B:
Real Money Stock in First Differences

Figure 2A:
Real GDP
Figure 2B:
Real GDP (in First Differences)

Figure 3A:
Treasury-Bill Rate

Figure 3B:
Treasury-Bill Rate (in First differences)
Figures 4A and 4B are graphs of the interest-rate spread (s) in levels and first-difference, respectively. Both show no discernible trend. As in the case of the treasury-bill rate, the interest-rate spread also remained relatively stable throughout the pre-reform period. It, however, spiked up soon after the reforms in 1986 and reached an all-time high in 1992. Thereafter, it fell swiftly and has maintained a zigzag pattern since. Its first-difference, on the other hand, stayed mostly within the 1.96 standard-error bands, except in 1990, 1992, and 1993.

The inflation rate also shows a slight positive trend initially up to 1987, as shown in Figure 5A. Thereafter, the trend became slightly negative and it remained so throughout the rest of the sample period. On the other hand, the first-difference of the inflation rate oscillated around its mean throughout the sample period. It fell outside the 1.96 standard-error bands only twice (in 1987 and 1988).
Finally, it is always advisable to plot the endogenous variables in a VAR model together in order to determine whether they are cointegrated and to look for clues about the type of deterministic terms to expect in their co-integrating relation(s). However, because real money balances and real GDP are in natural logs while the inflation rate and the interest rates are percentages, it is necessary to visually examine the two subsets of the data in separate graphs. Accordingly, real money balances and real GDP are presented in Figure 6A. They appear to be random-walk processes with drifts.

Figure 6B displays the Sierra Leone’s treasury-bill rate, the inflation rate, and the interest rate spread. As expected, there is comovement between them. Furthermore, the inflation rate exceeded the treasury-bill rate from the early 1970s through the early 1990s, even though interest rates had been liberalized in 1986. Thus, the real interest rate was negative throughout this period, suggesting a lack of sufficient incentives for economic agents to economize on their real money balances during this period.

On the basis of this visual examination, it can be concluded that there may be up to two cointegrating relations between the endogenous variables—one between the nominal variables (the domestic treasury-bill rate, the interest-rate spread, and the inflation rate) and another between real variables (the real money stock and real output). Finally, no visible trend is apparent in any of the hypothesized cointegrating relations.

**Figure 5A:**
**Inflation Rate**
Figure 5B:  
Inflation Rate (in First Differences)

Figure 6A:  
Money and Income (in 2000 Prices)

Figure 6B:  
Inflation and Interest Rates
IV. EMPIRICAL RESULTS

To apply the Johansen methodology, one must first determine the rank of the cointegrating matrix (II). Two alternative tests (the trace and maximum eigenvalue tests) have been routinely used for this. However, Cheung and Lai (1993), Haug (1996), and Johansen (2002) have shown that both tests are biased in small samples. To correct for this bias, Johansen (2002) developed a small-sample trace test that is based on the Bartlett-correction. This test has been incorporated into Version 2 of CATS in RATS. Additionally, by incorporating a multivariate stationarity test that is conditional on the deterministic terms in the model, Johansen’s procedure, as implemented in CATS in RATS, Version 2, dispenses with the need to conduct univariate unit-root tests before conducting the Johansen test.

Table 1:
Cointegration Rank Test

<table>
<thead>
<tr>
<th>Null Hypothesis</th>
<th>Eigenvalue</th>
<th>Uncorrected Trace</th>
<th>Corrected Trace</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Stat.</td>
<td>p-value</td>
<td>Stat.</td>
</tr>
<tr>
<td>r = 0</td>
<td>0.7615</td>
<td>0.0151</td>
<td>110.8070</td>
</tr>
<tr>
<td>r ≤ 1</td>
<td>0.4602</td>
<td>0.6405</td>
<td>57.1447</td>
</tr>
<tr>
<td>r ≤ 2</td>
<td>0.3671</td>
<td>0.8049</td>
<td>33.7628</td>
</tr>
<tr>
<td>r ≤ 3</td>
<td>0.2167</td>
<td>0.9058</td>
<td>15.9428</td>
</tr>
<tr>
<td>r ≤ 4</td>
<td>0.1421</td>
<td>0.8220</td>
<td>6.2370</td>
</tr>
</tbody>
</table>

Table 2:
Normalized Cointegrating Vector and Adjustment Parameters

<table>
<thead>
<tr>
<th>Vector</th>
<th>m_t</th>
<th>y_t</th>
<th>r_t</th>
<th>s_t</th>
<th>Δp_t</th>
<th>R</th>
<th>W</th>
</tr>
</thead>
<tbody>
<tr>
<td>β1</td>
<td>1.0000</td>
<td>-1.5194</td>
<td>0.0551</td>
<td>-0.0664</td>
<td>-2.2388</td>
<td>1.4662</td>
<td>-0.5229</td>
</tr>
<tr>
<td></td>
<td>(2.9535)</td>
<td>(1.9920)</td>
<td>(2.5862)</td>
<td>(6.8234)</td>
<td>(7.177)</td>
<td>(2.1650)</td>
<td></td>
</tr>
<tr>
<td>α</td>
<td>0.0211</td>
<td>0.0769</td>
<td>4.3806</td>
<td>3.8959</td>
<td>0.1969</td>
<td>--</td>
<td>--</td>
</tr>
<tr>
<td></td>
<td>(0.4537)</td>
<td>(4.3896)</td>
<td>(1.4679)</td>
<td>(1.2728)</td>
<td>(5.5165)</td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

a = Significance at the 0.01 level; b = Significance at the 0.05 level.

For these reasons, CATS in RATS, Version 2, was used to conduct the cointegration analysis in this paper. Based on the Schwarz Criterion, which
Monte-Carlo studies show to be the most robust of the available model-selection criteria (Bessler and Binkey 1982, Geweke and Meese 1981, and Oh 2002), the endogenous variables in the model were lagged once. Finally, to account for the impact of the war and the financial reforms of 1986, two dummy variables—for the war years (W) and the post-reform period (R), respectively—were included as deterministic terms.

The model was estimated with an unrestricted constant in the cointegrating relation. Table 1 reports trace statistics (both uncorrected and small-sample-corrected) for determining the rank of the cointegrating matrix ($\Pi$). The first column states the null hypotheses (that $r=i$ against the alternative that it exceeds $i$). The estimated eigenvalues are reported in the next column. Meanwhile, the third and fourth columns, respectively, report the uncorrected and corrected trace statistics and their probability values for the null hypothesis that the rank of $\Pi$ is equal to $i$.

Based on the evidence provided in Table 1, both tests reject the null hypotheses that $r=0$ (against the alternative that it is at least 1) at the 0.05 level. However, they fail to do so at all other levels. This suggests that $r=1$, which means that there is a unique long-run equilibrium relation between the variables.

The estimated cointegrating vector ($\beta'$) and the adjustment vector ($\alpha$) are reported in Table 2. The first row reports the estimated coefficients and the absolute values of their t-statistics. The t-tests reveal that each estimated coefficient is significantly different from zero at the 0.05 level or better.

Given the nature of financial intermediation in Sierra Leone, people generally hold their incomes in cash balances. Therefore, the money demand function should be unit income-elastic. This proposition is supported by the fact that the null hypothesis that $\beta_2 = 1$ could not be rejected.

Furthermore, as expected, the estimated interest-rate elasticity of money demand $\beta_3$ is negative. This means that Sierra Leoneans economize on real money balances when the opportunity cost of money rises, and vice versa. Also, as expected, the estimated coefficient of the interest-rate spread ($\beta_4$) is positive. This suggests that, all other things being equal, an increase in the interest-rate spread ($s$) makes domestic assets (including real money balances) more attractive to Sierra Leoneans. Moreover, although the estimates of $\beta_3$ and $\beta_4$ are different in absolute values, they are not significantly different from each other. This suggests that dollar- and leone-denominated-assets are perfect substitutes in the portfolios of economic agents in Sierra Leone.

The expected inflation rate was found to be positively related to the demand for real money balances in the long-run. Specifically, the evidence shows that an increase in the inflation rate by 1 percent leads to a 2.2388 percent increase in the long-run demand for real money balances. This suggests that the “investment” effect of inflation on money demand outweighs the “hedging” effect.
The estimated coefficient of the war dummy had its expected positive sign, thereby implying that the insecurity of the war years boosted the demand for liquidity by an average of 0.5229 percent. Finally, the estimated coefficient of the post-reform dummy was negative and significant. This suggests that by providing Sierra Leoneans with the option to hold foreign-currency-denominated assets and allowing the domestic interest rates to be market determined, the financial reforms of the mid-1980s caused a reduction in the mean desired money balances by 1.4662 percent.

The second set of results reports the estimated short-run adjustment parameters \( (\alpha) \) and the absolute values of their associated t-statistics. Except that they measure the speed of adjustment of the endogenous variables toward their long-run equilibrium values following a short-run shock, the elements of \( \alpha \) are of no economic significance. Thus, a finding that a given \( \alpha_i \) is not significantly different from zero (as in the case of \( \alpha_1, \alpha_3 \) and \( \alpha_4 \)) means the absence of an error-correcting component to the cointegrating relation for the corresponding variable. Accordingly, the variable is said to be weakly exogenous to the cointegrating relation (Johansen 1992).

For the cointegrating relation to be equilibrium-correcting in a given \( \Delta x_{it} \) equation, the signs of each significant \( \alpha_{ij} \) must be reversed. In other words, if a given significant \( \Delta_{ij} \) is positive, \( \beta_{ij} \) must be negative. Otherwise, the cointegrating relation would be characterized by overshooting (Juselius 2007: 122). Because \( \alpha_2 \) and \( \alpha_5 \) are positive while \( \beta_2 \) and \( \beta_5 \) are negative, Table 2 shows that the estimated cointegration relation is equilibrium-correcting in their associated variables.

V. DIAGNOSTIC TESTS

The strength of the above results depends on whether the underlying assumptions of the model are satisfied. Firstly, for a cointegrating relation to exist between variables, they must be non-stationary. A second crucial assumption is that the residuals must be normally, independently, and identically distributed (NIID). Moreover, the robustness of the estimated long-run money demand relation depends on the constancy of the estimated parameters and the stability of the model. Parameter constancy is particularly important for this study because Sierra Leone went through not only a change in economic and political regimes but also an eleven-year brutal civil war during the sample period.

A. VARIABLE STATIONARITY

Table 3 reports the results of a multivariate stationarity test for the endogenous variables in the model. They show that, conditional on the
rank of the co-integrating vector being 1, as determined earlier by the trace test, the null hypothesis that each of the endogenous variables is stationary can be rejected at better than the 0.01 level. This justifies the use of cointegration analysis to estimate the long-run equilibrium relationship between these variables.

**Table 3**

Multivariate Stationarity Tests

<table>
<thead>
<tr>
<th>Rank</th>
<th>d.o.f.</th>
<th>m</th>
<th>y</th>
<th>r_b</th>
<th>s</th>
<th>Δp</th>
</tr>
</thead>
<tbody>
<tr>
<td>1</td>
<td>4</td>
<td>45.4954a</td>
<td>36.6017a</td>
<td>27.1697a</td>
<td>28.0743a</td>
<td>17.7334a</td>
</tr>
<tr>
<td>2</td>
<td>3</td>
<td>12.5236b</td>
<td>14.4602a</td>
<td>5.6584</td>
<td>5.5906</td>
<td>8.6715b</td>
</tr>
<tr>
<td>3</td>
<td>2</td>
<td>7.6339b</td>
<td>7.9554b</td>
<td>2.6505</td>
<td>2.6431</td>
<td>7.6141b</td>
</tr>
<tr>
<td>4</td>
<td>1</td>
<td>0.1115</td>
<td>0.5674</td>
<td>0.3417</td>
<td>1.7173</td>
<td>2.3142</td>
</tr>
</tbody>
</table>

*a* = Significance at the 0.01 level; *b* = Significance at the 0.05 level.

**Table 4**

Residual Tests

<table>
<thead>
<tr>
<th>Test</th>
<th>Lag</th>
<th>X^2</th>
<th>d.o.f.</th>
<th>p-value</th>
</tr>
</thead>
<tbody>
<tr>
<td>LM</td>
<td>1</td>
<td>21.9735</td>
<td>25</td>
<td>0.6373</td>
</tr>
<tr>
<td></td>
<td>2</td>
<td>33.0598</td>
<td>25</td>
<td>0.1296</td>
</tr>
<tr>
<td>ARCH</td>
<td>1</td>
<td>250.7936</td>
<td>225</td>
<td>0.1144</td>
</tr>
<tr>
<td>Normality</td>
<td>--</td>
<td>13.0327</td>
<td>10</td>
<td>0.2218</td>
</tr>
</tbody>
</table>

B. Residual Analysis

To confirm that the residuals from the model are *NIID*, multivariate versions of three commonly used residual tests—the Breusch-Godfrey LM test (for serial correlation), the ARCH test (for heteroscedasticity), and the Jarqui-Bera test (for normality)—were conducted. The results, which are reported in Table 4, show that the residuals are normally distributed and not serially correlated at both the first and second lags. The ARCH test also fails to reject the null hypothesis of homoscedasticity at the first lag. These findings suggest that the residuals from the VAR are indeed *NIID*. 
C. Model Stability

The stability of a VAR model requires that its characteristic roots lie within a unit circle, meaning that the absolute value of each must be less than one. Table 5 shows that this condition is satisfied because all the characteristic roots are less than unity in absolute value.

Table 5
Roots of VAR(1) Model

<table>
<thead>
<tr>
<th>_i</th>
<th>Real</th>
<th>Imaginary</th>
<th>Modulus</th>
</tr>
</thead>
<tbody>
<tr>
<td>1</td>
<td>0.7093</td>
<td>0.0151</td>
<td>0.7095</td>
</tr>
<tr>
<td>2</td>
<td>0.7093</td>
<td>-0.0151</td>
<td>0.7095</td>
</tr>
<tr>
<td>3</td>
<td>0.3624</td>
<td>0.2335</td>
<td>0.4311</td>
</tr>
<tr>
<td>4</td>
<td>0.3624</td>
<td>-0.2335</td>
<td>0.4311</td>
</tr>
<tr>
<td>5</td>
<td>0.3420</td>
<td>0.0000</td>
<td>0.3430</td>
</tr>
</tbody>
</table>

D. Parameter Constancy

To test for parameter constancy, the model was estimated recursively using 1964-85 (pre-reform) as the baseline period. The recursively estimated beta vector for the post-baseline period (1986-2005) period was normalized such that a value of 1.0 became the 0.05 confidence limit for the null hypothesis that the baseline beta vector is not significantly different from its post-baseline counterpart. Thus, the null hypothesis of parameter constancy is rejected whenever the normalized values of the post-baseline beta reaches or exceeds unity.

The normalized beta for the post-baseline period is plotted in Figure 7. It starts at the 0.5 mark and declines steadily toward zero. Because it comfortably remained way below the 1.0 critical point, Figure 7 rejects the null hypothesis of parameter non-constancy in the baseline and post-baseline periods.

E. Predictive Ability

How well a model predicts the endogenous variables in a VAR system is important diagnostic test of its performance. To demonstrate this, actual and predicted values of the log of real money balances, the variable of interest here, are presented in Figure 8. It shows that the model does a reasonably good job of predicting real money balances, thus providing further support for its appropriateness to this analysis.9
The robustness of the model was further investigated by restricting the coefficients of the explanatory variables ($\beta_2$, $\beta_3$, $\beta_4$, and $\beta_5$) to zero. A failure to reject this restriction would mean that all the endogenous variables in the money demand relation are strongly exogenous to the cointegrating relation. The chi-square statistic for testing this hypothesis amounted to 32.7095. With 4 degrees of freedom, the null hypothesis was rejected at better than the 0.0001 level. This means that past values of the explanatory variables in the money demand equation do indeed explain variations in Sierra Leone’s equilibrium real money stock.

VI. SUMMARY AND CONCLUSIONS

A key motivation for this study was to use current advances in time-series econometrics to estimate Sierra Leone’s real money demand function for the 1964-2005 period and compare the results to those reported in Kallon (1992). The two studies are similar in the sense that the long-run
interest-rate and income elasticities of money demand have their predicted signs. However, their absolute values are larger in the current study than in Kallon (1992). This can be explained by differences in model specification, structural changes introduced by the financial liberalizations of the mid-1980s and the civil war (both of which were absent in the earlier study), and differences in the frequency of the data used in the two studies. Specifically, whereas this study uses annual data, Kallon (1992) used quarterly data. And as economic theory teaches, the absolute value of the elasticity of demand for a product is larger when economic agents have a longer time to adjust to changes in the relevant variable.

Thus, the absolute value of the long-run interest elasticity reported in Kallon (1992), 0.0265, is much smaller in absolute value than the 0.0980 reported in this study. Obviously, allowing Sierra Leoneans to hold foreign-currency-denominated assets must have also made the demand for money more interest elastic in the post-1986 period. This supports the Gurley-Shaw (1955, 1960, 1967) hypothesis that, by providing alternative assets to economic agents, improved financial intermediation makes the demand for money more interest elastic. Even so, the interest-rate elasticity of money demand is very small in both studies, which is expected in economies with under-developed financial markets.

As expected in economies where people hold their non-real assets in money balances, the estimated income elasticity of long-run money demand in Sierra Leone is not significantly different from unity even though its numerical value is 1.5194. This suggests that there are no economies of scale in money-holding in Sierra Leone.

Policy-wise, the relatively low interest rate elasticity of money demand suggests that monetary policy can be an effective short-term stabilization tool for the country. The underlying assumptions for this conclusion are: (1) that banks lend out all their excess reserves; and (2) that there is a ready-demand for them. However, the structural characteristics of Sierra Leone’s financial system belie these assumptions. Firstly, a vast majority of Sierra Leoneans have no access to bank services. Therefore, commercial banks have no information on their creditworthiness. Consequently, as demonstrated by interviews I conducted in 2002 with managers of the five commercial banks operating in the country at the time, banks in Sierra Leone actually prefer holding excess reserves (which they use to buy government debt obligations) to lending to customers whose credit-worthiness cannot be determined. Thus, a monetary expansion might not filter down to potential investors in especially the rural areas.

Secondly, most payments in the country are made in cash, which is held by the public. This cash drain on the banking sector causes the money multiplier to be smaller than its theoretical maximum (which is the inverse of the required reserve ratio). This makes the impact of monetary policy on the real
sector very small—even though money demand is very interest-rate inelastic.

Thus, policies that would ameliorate this problem and thus enhance the
efficacy of monetary policy are sorely needed in Sierra Leone. Accordingly,
this study recommends bringing modern financial institutions and/or micro-
finance facilities to rural Sierra Leoneans. As McKinnon (1973) and Shaw
(1973) have shown, the resultant financial deepening from such a move will
augur well for capital accumulation and long-run economic growth.

Perhaps now is the time for the monetary authorities to transform the
currently dormant Post Office Savings Bank into micro-credit institutions
and locate at least one per chiefdom. With the specialized knowledge of
local conditions that they will eventually acquire, these institutions will be
able to profitably provide financial services to the nation’s rural population
which currently does not have access to such services. This will create
a mechanism through which rural Sierra Leoneans can have access to the
excess reserves of the banking system, thus enhancing the effectiveness of
monetary policy as a short-run stabilization tool. Finally, the promotion of
the banking habit among rural Sierra Leoneans, the mobilization of rural
savings into the financial system, and the removal of self-finance as the key
vehicle for financing investment projects by most Sierra Leoneans are the
expected additional benefits of such a policy.

Notes

1. For more on the importance of self-finance on entrepreneurial activity
in Sierra Leone, see Kallon (1990: 185 – 191).
2. With the 1971 values of 36.0 and 0.224223, respectively, for the 1980
and 2000 base years, the 1964-1970 GDP deflator (1980 = 100) was trans-
formed to the 2000 base-year by multiplying it by 0.006228.
3. The estimated regression equation is as follows:
   \[ x_t = \beta_1 m_t + \beta_2 y_t + \beta_3 r_t + \epsilon_t \]
   where \( x \) = natural log of the nominal (leone-dollar) exchange-rate, \( t \) =
trend term for post-reform period), \( D \) = for the post-reform period; zero
otherwise, and \( \epsilon_t \) is the \( i \) th order moving-average term. The t-statistic for
each coefficient is in parenthesis (the supercript \( a \) and \( b \) denotes significance
at the 0.01 and 0.05 levels, respectively). A shift dummy for the war years
(\( W = 1 \) for 1991 – 2001; zero otherwise) had no significant impact on the
dependent variable.
4. For a thorough exposition of the Johansen cointergrated VAR model
and how to estimate its parameters, see Juselius (2006).
6. The long-run equilibrium condition is as follows:
   \[ \beta_1 m_t - \beta_2 y_t - \beta_3 r_t - \epsilon_t \]
\[ \beta_4 \delta t - \beta_5 \Delta p_t = 0. \] When normalized in terms of \( m \), \( \beta_1 \) is assumed to be equal to 1.

7. The chi-square statistic for testing this hypothesis equaled 0.4667. With one degree of freedom, it was too small to justify the null hypothesis.

8. With one degree of freedom, the computed chi-square statistic of 1.1113 is not large enough to reject the null hypothesis that \( \beta_2 = \beta_3 \) in absolute value.

9. The model predicted the other endogenous variables very well. However, in the interest of brevity, they are not reported here.

**References**


______ (1992). “Testing Structural Hypothesis in a Multivariate Cointe-
gration Analysis of PPP and the UIP for UK.” *Journal of Econometrics* 53: 211 – 244.


