The demand for money in Jamaica: a cointegration approach

I. INTRODUCTION

The hypothesised existence of a well-defined money demand function constitutes an important part of the philosophical content of conventional monetary analysis. However, researchers in the Caribbean (for example, Bourne [2], Howard [31], and Ramsaran and Maraj [48], have hitherto failed to provide satisfactory evidence of a highly stable demand for money function in any of the territories examined. Ordinary least square estimates from these studies are plagued with serial correlation in the residuals and as Granger and Newbold [22] point out, serial correlation constitutes evidence of misspecification. Moreover, most of these studies have followed the ad hoc practice of including as one of the regressors in the money demand function, a lagged dependent variable to deal with the problem of incomplete adjustment of money demand in the short run. Such action is justified by the partial adjustment mechanism (Nerlove [45]) or the adaptive expectations hypothesis propounded by Cagan [5]. However, it is well known that the partial adjustment mechanism constrains the adjustment pattern in the

Mr. Craigwell is an Economist from the Central Bank of Barbados. He wrote the paper at the University of Warwick and would like to thank Kenneth Wallis, Mark Stewart and Anthony Wood for helpful comments.

1 This literature is surveyed in Farrell and Christopher [17].
regressand to be the same regardless of the source of the initial disturbance, whilst the adaptive expectations model, as noted by Muth [44], is only optimal when the data generating process follows an autoregressive integrated moving average process (ARIMA) of degree (0,1,1).

Recently, Hendry and Mizon [30] and Hendry [27], showed that ‘error correction’ mechanisms (ECM), which encompass models in both levels and differences, and are compatible with proportional long-run equilibrium behaviour, have less restrictive lag structures and nest the partial adjustment and adaptive expectations models. The ECM therefore circumvent the fundamental ‘spurious’ regression problem discussed by Granger and Newbold [22] through the use of appropriate difference variables in the model, but without losing long-run information due to using differenced data only. More support for the use of ECM comes from the recently developed theory of cointegration. This theory allows estimation and inference to be possible when economic variables do not satisfy the classical assumptions of constant mean and variance. If there exists a linear combination of these non-stationary variables that is stationary, then these variables are said to be cointegrable. Engle and Granger [14] proved that if a set of variables is cointegrated then there exists a corresponding error correction form of those variables.

In this paper the demand for money function in Jamaica is examined using cointegration theory and the error correction mechanism. Section II of this paper discusses the specification of the demand for money function, while in Section III the econometric methodology is presented. The data and empirical results are given in Section IV, followed finally by some conclusions. Heavy emphasis is placed on diagnostic tests as a means of evaluating model adequacy prior to the examination of the theoretical prediction. Special attention is also drawn to the role of the encompassing principle for model evaluation since an acceptable approximation to the data generation process should be able to explain contending findings.

II. THE ‘LONG-RUN’ DEMAND FOR MONEY FUNCTION

As the monetary sector in less developed countries (LDC) tends to be underdeveloped and financial alternatives to money are few, most of the work done on the demand for money in LDC has focussed on the correct composition of the vector of opportunity costs. Researchers have generally used the rate of (expected) inflation as a proxy for the opportunity cost on real assets and a representative (vector) of domestic interest rates to reflect the cost on domestic financial assets. It is usually assumed that observed domestic rates are institutionally pegged and consequently, may not adequately reflect

\[2\] This point and other criticisms of the partial adjustment mechanism can be found in Laidler [37].
domestic money market conditions. The holding of real assets such as land, jewelry and houses may in fact provide the only meaningful alternative to holding money. But as data on the actual rates of return on real assets are virtually non-existent in the Caribbean, it is argued that one could proxy these rates by the rate of inflation or the expected rate of inflation. Of course, inflation expectations can also enter into money demand functions on its own merit, since in a world of embryonic and highly imperfect capital markets, observed interest rates may not fully reflect inflation expectations (see Fama [16]).

The openness and institutional setting of a country can also influence the form and content of the money demand function. In particular, where there is a (free) flow of funds between an underdeveloped economy and the rest of the world, foreign interest rates theoretically affect the domestic demand for money (Hamburger [25]). In effect foreign interest rates capture the degree of substitutability between local and foreign assets.

In this study, the foreign interest rate, neglected so far by Caribbean economists, is included in the money demand function, along with a representative domestic interest rate and the rate of inflation, to capture the opportunity cost aspect of the demand for money. The full ‘long-run’ demand for money function is specified mathematically as follows:

\[ M_t = C_0 + C_1 P_t + C_2 Y_t + C_3 R_t + C_4 \text{INF}_t + C_5 \text{FR}_t \]

where \( M_t \) is the nominal demand for money, \( P \) is the price level, \( Y \) is the real constraint variable, \( R \) is the domestic interest rate, \( \text{INF}_t \) is the rate of inflation while \( \text{FR} \) is the foreign rate, \( C_0 \) is a scale factor and \( C_i, i=1...5 \), are the variables (in logs) corresponding to long-run elasticities, \( C_1 \) and \( C_2 \) are expected to be positive while \( C_3 \), \( C_4 \) and \( C_5 \) are expected to be negative. The choice of the log linear form recognises that there may be some interaction between the effects of the regressors, and moreover, it allows for comparisons with studies on the demand for money which we will refer to in this paper.

With regard to the scale variable, we follow the view that the demand for money is best regarded as constrained by transactions. The majority of studies which adopt this approach use income, or some expectational variant thereof, as a proxy for transactions. In this paper an “augmented transaction proxy”-sum of income and imports-suggested by McClean [43] is also employed to allow for the possibility that in a small open economy undergoing

---

3 Wong [53] argues that if there are perceived linkages between organised and non-organised credit markets and if borrowing is a major source of financing economic activity, interest rates in the unorganised markets, although unobservable, would tend to mirror the degree of credit constraint in the economy. Caribbean economies, however, although characterised by relatively underdeveloped money and capital markets, tend to be highly monetised and informal financial arrangements may be of little account (Worrell and Prescod [55]).

4 This is not a universal view; see for example, Laidler [36].
significant structural change, and characterised by a high level of foreign trade transactions which dominate monetary movements, income may not be a reliable proxy for transactions.\(^5\)

Another feature of conventional money demand analysis is to use the real stock of money as the dependent variable in regression analysis. One rationale for this is the assumption that the nominal stock of money is best regarded as an exogenous variable, for example, see Friedman [18]. Secondly, the nominal demand for money is considered to be homogenous of degree one in prices, thus restricting the price elasticity of the demand for money to one.

However, in LDC, the high degree of openness of these economies, suggests that the nominal supply of money may be an endogenous variable and can therefore be employed legitimately as the regressand (for an elaboration of this point see McClean [42]). In addition, McClean [43] argues that "the linear homogeneity postulate has not been subjected to extensive testing, and there is strong circumstantial evidence to suggest it should be. For example, money demand models which restrict the price elasticity of demand to unity are usually plagued by serial correlation. Given the pronounced time trend in the price level, a price elasticity of demand for money which is significantly different from unity could explain the presence of serial correlation in such models". For these reasons we do not follow previous money demand studies in the Caribbean (eg. Bourne [2]) and impose the linear homogeneity restriction a priori, instead we perform tests of its validity. These tests are carried out in Section IV, but before we present the results, let us discuss the econometric methodology used in this paper.

III. THE ECONOMETRIC METHODOLOGY\(^6\)

Traditional econometric theory has been developed on the postulate that the underlying data processes are stationary, despite the fact that most economic variables exhibit non-constant mean and variance. In these circumstances classical inference is rendered invalid, as evidenced by the substantial literature on 'spurious regressions'. However, new developments in econometric theory has shown that valid estimation and inference is possible, if there exists a linear combination of these non-stationary variables that is stationary. If a linear combination of non-stationary economic series is stationary, these series are said to be cointegrated and can be regarded as defining a 'long-run equilibrium' subspace. Thus these series should not diverge from each other, at least in the long run, by too great an extent. The

---

\(^5\) Other reasons why income may not be a reliable proxy for transactions are well documented in Lieberman [39].

\(^6\) For a thorough discussion of this methodology see Leon [38].
variables may drift apart in the short run but in the long run they are brought together by economic forces such as a market mechanism or government intervention. Cointegration theory therefore permits the separation of the long-run information contained in the data from the complex dynamics, about which economic theory is generally silent. Cointegrating variables can be defined formally as follows: if we call a stationary series after differencing $d$ times to be integrable of order $d$, $I(d)$, then a set of variables, $Z_t$, are cointegrated of order $(d, b)$, $CI(d, b)$, if there exists a cointegrating vector $a' \neq 0$ such that $e_t = a'Z_t \sim I(d, b), \quad b > 0$ (see Engle and Granger [14]). Where the dimension of $Z_t > 2$, vector $a$, need not be unique as there may be several cointegrating vectors or equilibrium relationships, some of which may be linearly dependent.

The methodology adopted here follows the above theoretical developments. Specifically, we utilise the Granger Representation Theorem (see Engle and Granger [14]) which states that, if a set of variables cointegrate, then there exist corresponding ECM of those variables. More formally, for $r$ linearly cointegrating vectors and $d = b = 1$, there exists an error correction representation of $r$ stationary random variables, such that

\[ (2) \quad A^*(L)(I-L)Z_t = be_{t,1} + g(L)U_t, \]

where $A^*(o) = I_n$, $N$ is the dimensionality of $Z_t$, and $g(L)$ is a finite scalar lag polynomial. The Granger Representation Theorem therefore provides a sound theoretical basis for ECM when the level terms cointegrate. Furthermore, if the data generating process (DGP) is an equation like (2) then $Z_t$ must be a set of cointegrating variables. The practical implications for dynamic modelling are profound.

Engle and Granger [14] show that if OLS is used to estimate the cointegrating vector then the other parameters of the ECM may be consistently estimated provided that the first stage estimates of the cointegrating vector are imposed on second stage ECM. This is done by including the lagged error terms from the cointegrating regression in a general ECM. This procedure is sometimes referred to as the Granger-Engle two-step procedure. This method gives ‘second stage’ OLS standard errors that are consistent estimates of the true standard errors. The two-step procedure uses super convergence properties of the first stage estimates and permits one to test whether the

---

7 This theorem is not wholly uncontroversial; Wickens and Breusch [52] argue that there is not a strong relationship between cointegration theory and the conventional ECM but Granger and Engle’s version of the ECM can be interpreted as just another transformation of the original dynamic model.

8 If Wickens and Breusch’s comments are taken to be valid, then this two-step method will be unnecessary and estimation is done by estimating the short and long-run parameters simultaneously using the instrumental variables method.
vector of variables properly cointegrates or not. Thus the full ECM is not a spurious regression.

Since testing for a cointegrated vector requires the series to be integrable of order $d>0$, we proceed in the following way: (i) investigate the temporal properties of the variables in the static 'long-run' equilibrium formulation of interest; (ii) test for a vector of cointegrated variables; (iii) estimate the ECM, and (iv) test the adequacy of the resulting equation using standard diagnostic tests and encompassing principles.

IV. DATA AND EMPIRICAL RESULTS

The results we discuss next were estimated with Jamaican data that span the period 1953-1986, using the following definitions of the variables:

The money stock was defined as end of period currency held by the public plus end of period demand deposits at commercial banks. In the Caribbean, writers such as McClean [41] and Prescod and Worrell [55], have argued that time deposits are not considered by individuals to be important relative to the narrow concept of money. In particular, agents consider other assets, mainly in the form of jewelry, as the preferred alternative to narrow cash balances. Such holdings are however impossible to measure, and inclusion of time deposits in the money stock lead to definitional problems. The statistical data of the narrow money concept used in this paper is taken from International Monetary Fund (IMF), International Financial Statistics (IFS), line 34.

The income aggregate employed in this study is gross domestic product (GDP) at market prices, which is based on the national income accounts. Problems with such series are well known and stem primarily from changes in the methodology used to assemble the accounting data, and to potentially large errors in the survey data. Nevertheless, these are the best income data available at this time. Some writers have used gross domestic expenditure as a possible alternative to gross domestic product, but recent work by Domowitz and Hakkio [11] revealed that differences in the estimated relationships are minor and not statistically significant. The GDP data is obtained from line 99b of the IFS.

The only price series available over the whole sample period are the consumer price index (CPI), taken from line 64 of the IFS. This composite index, whose base period is January 1975, evolved from data obtained in a Household Budget Survey conducted in 1971-72. The indices relate to lower and middle income households whose total household expenditure in 1971-72 was 4000 dollars or less, representing 85% of all households. Since 1975 the data have been continually reweighted, leading to different base years. The series are linked by using ratio splicing at the first annual overlap and the

---

9 The 640K version of DATAFIT computer program was used throughout this exercise.
linked series are shifted to a common base period, in this case 1980 = 100. This index is the deflator used for all real variables referred to in the text.

The annually averaged (average of monthly data) treasury bill rate was used as a representative rate for the domestic economy. No other domestic rate of interest was available for a sufficiently long period. The comparable foreign interest rate is the annual average of end of period weekly data treasury bill rate of the United Kingdom. These rates are found in IFS, line 60C.

Finally, the import variable, used in the 'augmented transactions proxy' is defined as the c.i.f. value of imports. The source is IFS, line 71.

Estimation

From the procedure outlined above in section III we first determined the order of integration of the individual series in equation (1). In this regard we employ tests for unit roots suggested by Dickey and Fuller [8,9]. The Dickey and Fuller tests are based on the \( t \) value of \( Z_{p,i} \) in the following OLS regression:

\[
(3) \quad DZ_i = d_0 Z_{p,i} + \sum_{j=1}^{p} d_j DZ_{p,j} + e_i,
\]

where \( P \) is chosen to be sufficiently large to ensure that the residual \( e_i \) is empirical white noise. When \( P=0 \) the Dickey Fuller (DF) test is defined and \( P=0 \) specifies the Augmented Dickey-Fuller (ADF) test. The null hypothesis that \( Z_i \) follows a random walk is rejected if \( d_0 < 0 \) and significantly different from zero.

The above testing methods are applied to the logarithms of the Jamaican time series over a reduced estimation period (1958-1986), to accommodate the estimation of two lagged dependent variables \( (P = 0, 1, 2) \) after third order differencing. The use of a common estimation period facilitates comparisons of alternative lag truncation. Since the Dickey and Fuller tests may lose power when the identical independent distribution assumption is not valid (see Phillips [47] for further discussion) the residuals are tested for serial independence using a Lagrange multiplier test suggested by Pagan and Hall [46], and for heteroscedasticity in the form of a Pagan and Hall's [45,p.178] suggested alternative to White's [51] test. The results of the unit root tests indicate that all the variables are I(1) [see Table 1] implying that they need to be differenced once to transform them to stationary series.

Next we investigate whether equation (1) is a cointegrated set or not. If the error term of this appropriately static regression is stationary, then the specified vector of variables will form a cointegrated set. Estimating equation (1) over the period 1954 - 1986 we obtain the following OLS logarithmic regression results:

\[10\] The formulation used for the latter test includes the fitted values up to the second power.

\[11\] The results presented here and afterwards use real income rather than the 'augmented transaction proxy' of McClean, since the results of the latter model with it included were not cointegrated (DF = -
### Table 1. Jamaica: Unit Root Statistics

<table>
<thead>
<tr>
<th>Variable</th>
<th>DF</th>
<th>ADF</th>
</tr>
</thead>
<tbody>
<tr>
<td>Narrow money</td>
<td>-5.23</td>
<td>-2.21</td>
</tr>
<tr>
<td>Real income</td>
<td>-3.81</td>
<td>-3.47</td>
</tr>
<tr>
<td>Domestic interest rate</td>
<td>-5.21</td>
<td>-4.05</td>
</tr>
<tr>
<td>Foreign interest rate</td>
<td>-4.89</td>
<td>-5.79</td>
</tr>
<tr>
<td>Prices</td>
<td>-2.24</td>
<td>-2.56</td>
</tr>
</tbody>
</table>

*The critical value used is the 5% level (-1.95). See Fuller [19, p.373].

\[
M = -2.826 + 1.257P + 1.098Y - 0.231R - 0.278INF - 0.080FR \\
R^2 = 0.997 \quad S.E.R. = 0.075 \quad DW = 1.789 \quad DF = -5.420 \\
(ACF)LAG 1-11: 0.095 0.057 0.162 0.028 -0.063 0.041 -0.444 -0.292 0.135 -0.167 -0.109
\]

Standard errors of coefficients are not reported for the estimated cointegration parameter since test statistics are not well developed. Further, interpretation of the magnitude of the cointegration parameter can be problematic. Benerjee et al. [1] report that the small sample bias in estimating the cointegration parameter diminishes as the \( R^2 \) value approaches one. Since the reported cointegration regression has an \( R^2 \) value of 0.997, this may indicate that the bias is small in our case. The magnitude of the coefficient on \( P \) suggests that a real money demand function may be appropriate, but such an equation is not cointegrated (\( DF = -4.185 \) compared to a critical value of 4.76 at the 5% level). Hence we maintain the nominal specification.

The DF test statistic reveals that the residuals from equation (4) are \( I(0) \), indicating that the set of variables form a cointegrating set.\(^{12}\) The partial and autocorrelation function (ACF) of the residual vector also points to a stationary series, the Box-Pierce [3] and Ljung-Box [40] statistics for 11 lags being equal to 12.708 and 17.309 respectively compared to a 5% critical value of 19.675.\(^{13}\) There is however a spike in the residual correlogram at the seventh order, which was not common to this equation alone. This is not statistically different

---

\(^{4.033} \) Attempts at finding a long-run equilibrium subspace with the other four variables (when McClean's proxy was dropped) failed, suggesting that real income is a necessary variable for cointegrability to be achieved.

\(^{12} \) No critical values are available for six variables but Engle and Yoo [15] have calculated critical values up to five variables using Monte Carlo simulation techniques. These figures are used as an approximate guide in determining whether equations are cointegrated or not. For six variables we looked for any value significantly above five.

\(^{13} \) The Ljung-Box statistic is likely to perform better in small samples, see, for example Harvey [26, 11] for more details.
from zero at conventional significance levels but there seems to be no economic rationale for such a seventh order lag structure in this data and it may be reasonable to conclude that the observation is a chance occurrence which does not require remedial attention.

On the evidence above it is reasonable to conclude that we have a cointegrated set of variables and therefore OLS estimates that are superconsistent in the sense that they will converge more rapidly to their true value than the normal OLS estimates (see Stock [50]). Before going on to look at the error correction formulation there is a further complication which needs to be considered. Equation (4), although a valid cointegration regression, may not be unique. It is possible that any of the five regressors in equation (4) could have been used as the regressand in a regression. However, given the properties of OLS, the long-run equilibrium relation that results from the regression would not, in most cases, be identical to equation (4). It is therefore important to know just how many cointegrating vectors there are and how different the implied equilibrium subspace given by the different inversions of (4) would be. These questions are examined in Table 2 where the different inversions of equation (4) are depicted, along with some diagnostic statistics. As would be expected, the different inversions produce different estimates of the equilibrium parameters but the rank of the cointegrating matrix is clearly one - the equation with money demand as the dependent variable is the only one to obtain stationary residuals. Hall's [23] suggestion to use the equation with the highest $R^2$ would also point to the use of (4) for continued estimation, although there is a marginal difference between the $R^2$ of the price and of the money demand equations.

**Table 2. Jamaica: Cointegrability Results of the Money Demand Function**

<table>
<thead>
<tr>
<th>Inversions</th>
<th>$R^2$</th>
<th>$D$</th>
<th>$W$</th>
<th>$DF$</th>
</tr>
</thead>
<tbody>
<tr>
<td>$M=-2.83+1.26P+1.10Y-0.23R-0.28INF-0.08FR$</td>
<td>0.9972</td>
<td>1.79</td>
<td>-5.420</td>
<td></td>
</tr>
<tr>
<td>$P=2.23+0.77M-0.86Y+0.23R-0.19INF+0.07FR$</td>
<td>0.9971</td>
<td>1.76</td>
<td>-4.932</td>
<td></td>
</tr>
<tr>
<td>$Y=2.63-1.05P+0.82M+0.20R-0.17INF+0.11FR$</td>
<td>0.958</td>
<td>1.53</td>
<td>-4.412</td>
<td></td>
</tr>
<tr>
<td>$R=3.32+1.22Y+1.65P-1.04M+0.48INF+0.01FR$</td>
<td>0.933</td>
<td>1.19</td>
<td>-3.723</td>
<td></td>
</tr>
<tr>
<td>$INF=0.15+0.06R-0.13Y-0.17P+0.16M+0.06FR$</td>
<td>0.647</td>
<td>1.52</td>
<td>-4.421*</td>
<td></td>
</tr>
<tr>
<td>$FR=-3.41+1.01INF+0.03R+1.41Y+1.05P-0.74M$</td>
<td>0.813</td>
<td>1.10</td>
<td>-3.932*</td>
<td></td>
</tr>
</tbody>
</table>

(*) indicates serial correlation in the $DF$ formulation.

14 A more formal test is that of Johansen [34], who derives maximum likelihood estimates for all the cointegrating vectors in a given set of variables and a likelihood ratio test for the hypothesis that there is a certain number of cointegrating vectors. However, the limited degrees of freedom in our study make this procedure very difficult to implement as it requires convergence of maximum likelihood estimates.
Having achieved a suitable specification of the cointegration equation, we can proceed to the second stage of the Granger-Engle procedure. Following the 'general-to-specific' modelling methodology (see for example Hendry and Mizon [30]), an initially over-parameterised model with one lag on the dependent and independent variables was continually simplified and reparameterised until a parsimonious representation of the data generation process was obtained.\(^{15}\) The resulting equation for the period 1955-1984 is:

\[
(5) \quad DM = 0.036 + 1.344DP + 0.860DY - 0.151DR
\]

\[
(0.029) \quad (0.257) \quad (0.257) \quad (0.094)
\]

\[
-0.112DFR - 0.333DM(-1) - 0.645ECMT(-1);
\]

\[
(0.054) \quad (0.196) \quad (0.280)
\]

\[
R^2 = 0.605 \quad S.E.R. = 0.067 \quad DW = 1.911 \quad LMN[X^2(2)] = 4.072
\]

\[
HT1[F(1, 28)] = 0.541 \quad HT2[X^2(1)] = 0.042 \quad SCI[F(1, 22)] = 0.005
\]

\[
SC2[X^2(10)] = 10.576 \quad RESET[F(3, 26)] = 1.942 \quad WUT[F(3, 20)] = 0.215
\]

\[
CHOW[F(7, 18)] = 1.125 \quad PCI[F(2, 23)] = 1.143;
\]

where \(D\) is the first difference operator; \(ECMT(-1)\) is the error correction term from (4) lagged; \(LMN\) is the Jarque and Bera [33] normality test; \(HT1\) and \(HT2\) are a variant of White [51] and Engle’s [13] \(ARCH\) heteroscedasticity test respectively; \(SCI\) and \(SC2\) are a Lagrange multiplier test for first order serial correlation and Ljung-Box [40] statistic for lags of order 10, respectively; \(RESET\) is Ramsey’s [49] specification error test; \(WUT\) is Wu’s [56] \(T^2\) statistic for testing the independence of the regressors; \(CHOW\) is Chow’s [6] test for structural change or stability and \(PCI\) is a predictive accuracy test (see Chow [6]). Under the null hypothesis \(RESET, HT1, SCI, WUT\) and the two Chow tests have \(F\) distributions with degrees of freedom as indicated in the parentheses, while \(LMN, HT2\) and \(SC2\) follow chi square distributions with 2, 1, and 10 degrees of freedom, respectively. The \(F\) versions of \(SCI, HT1\) and \(PCI\) are presented here in the light of Kiviet [35] Monte Carlo evidence that these tests are more powerful in small samples. However, the \(X^2\) versions did not alter any of the underlying results. The numbers in brackets under the estimated coefficients are standard errors. The calculated \(F\) statistic in moving from the general to the parsimonious equation is 0.653, well below the tabulated value of 2.81, thus suggesting that equation (5) is an acceptable representation of the unrestricted form. Given this the model will now be evaluated in more detail.

The \(LM\) test for autocorrelation of the residuals is not significant, satisfying a necessary condition for white noise residuals. The Ljung-Box statistic is also supportive of a serial independence result. The implied marginal significance level for this latter test is roughly 0.40.

\(^{15}\) Due to multicollinearity between \(DP, DP(-1)\) and \(D^2P\), the term \(DP(-1)\) was dropped from the unrestricted regression.
The variant of White test fails to reject the null hypothesis of constant variance at conventional levels. The White test for general forms of heteroscedasticity is based on the auxiliary regression of the squared residuals on the elements of the upper triangle of the regressor cross-product matrix. But the test presented here is run on a subset of regressors to accommodate the small sample size and hence, circumvent the degree of freedom problem. As a result some loss of power can be expected. However, examination of the residuals with the Engle's ARCH test of first order reached a similar conclusion of homoscedasticity. Moreover, White's standard errors (not reported) were not noticeably different from those traditionally computed.

The correlation matrix (Table 3) reveals that the explanatory variables are nearly orthogonal variables, suggesting that multicollinearity is not problematic. Jarque and Bera test statistic indicates that the residuals are normal but the implied marginal significance of the test is about 0.17. Further investigation however showed that the moments of the scaled residuals (skewness 0.041 and kurtosis 3.089) are not significantly different from those of a standard normal distribution. The fitted regressions appear to plot the data quite well although there are some periods where there is a tendency to positive shocks (see figure 1). Indeed, the degree of fitness ($R^2=0.601$) is not extremely high for time series data but a high $R^2$ is not, however, a precondition for a well specified model, an interesting example is provided by Hendry ([27], equations 4 and 20). A more illuminating test of fitness is provided below using out of sample forecasting performance as the criteria. A final data coherence check was the RESET test for general functional form. This test has power against many forms of misspecification but is particularly useful when the maintained model has under-represented the curvature of the function it intends to estimate. The RESET test in this paper, following various simulation studies, uses polynomial terms of up to the fourth order, formed from the predictions of dependent variables from the maintained model. These added coefficients were rejected by the usual $F$ test for joint significance.

FIGURE 1. NARROW MONEY: ACTUAL & FITTED VALUES

--- DLNM --- DLNMF
TABLE 3. JAMAICA: ESTIMATED CORRELATION MATRIX OF VARIABLES FOR EQUATION (5)

<table>
<thead>
<tr>
<th></th>
<th>DY</th>
<th>DR</th>
<th>DP</th>
<th>DFR</th>
<th>DM(-1)</th>
<th>ECMT(-1)</th>
</tr>
</thead>
<tbody>
<tr>
<td>1</td>
<td></td>
<td>1</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>-0.008</td>
<td>1</td>
<td>0.298</td>
<td>1</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>-0.652</td>
<td></td>
<td>0.308</td>
<td>0.117</td>
<td>1</td>
<td></td>
<td></td>
</tr>
<tr>
<td>0.036</td>
<td>0.044</td>
<td></td>
<td>0.573</td>
<td>0.069</td>
<td>1</td>
<td></td>
</tr>
<tr>
<td>-0.228</td>
<td>-0.160</td>
<td>0.306</td>
<td>0.154</td>
<td>0.4932</td>
<td>1</td>
<td></td>
</tr>
<tr>
<td>-0.168</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

The stability of the model was checked over the period 1973-86 using the Chow test. This period is a particularly difficult period for Jamaica, relative to its past as there was a general slowdown in economic growth and relapses in the economy as a whole (see Worrell [54]). The Chow test statistic (PC2) is 1.125, somewhat below its critical 5% value of 2.58. Therefore the model appears to be stable. Note if the regression disturbances differ in variances in the two subsamples the Chow test result may be unreliable. An alternative test (PCI), also due to Chow [6], revealed for shorter periods, 1985-1986 and 1980-1986, that the model has stable coefficients. A less formal test, the cusum square plot of Brown et al [4], also suggests no evidence of significant parameter changes as it does not break its 95% confidence interval (see figure 2). This latter test is based on the cumulative sum of recursive residuals and is equivalent to a series of Chow tests, that is, Chow first test, PC2.

In a macroeconomic context, it is possible that price, income, and the domestic rate of interest are determined jointly with the demand for money. If this is the case, OLS estimates in the model will be inconsistent. The Wu-Hausman statistic (WUT) is used to test the independence between the regressors and the disturbance term. Based on this procedure, the null

FIGURE 2. CUMULATIVE SUM SQUARES OF RECURSIVE RESIDUALS

The straight lines represent critical bounds at 5% significance level.

16 A subset of instruments DR, DP and DY, each corresponding to regressors in the model, rather than instruments for all regressors is chosen due to lack of degrees of freedom. The instrumental variables used are $R_{r,1}$, $R_{r,2}$, $Y_{r,1}$, $Y_{r,2}$, $P_{r,1}$, and $P_{r,2}$. 
hypothesis of ‘exogeneity’ cannot be rejected at the 5% level of significance. Although they are indirect evaluations, the Chow and cusum square statistics also test weak exogeneity and, thereby valid conditioning, which is not rejected by the data.

Based on the battery of diagnostic tests, we are prepared to accept equation (5) as a well specified model in the statistical sense. This implies that the estimated model parameters can be given their proper economic interpretation. Nearly all the estimated coefficients are significant at the 5% level and of plausible size. The signs of the coefficients are those postulated in the model specification section. Growth in real income and prices increases the demand for money while changes in foreign and domestic interest rates (the latter significant at about the 15% level) reduces the growth in the demand for money. The lagged money term, statistically significant at approximately the 10% level, indicates that last period money stock impacts negatively on the current stock of money, suggesting some expectational adjustment mechanism in the “short run”. The error correction term is statistically significant supporting the previous conclusion that the variables are cointegrated. Further, it is noted that although the rate of inflation forms part of the long-run equilibrium relationship, changes in the inflation rate do not seem to be significant in explaining short-run changes in money demand.

Noteworthy is that the results support the claim that income effects on cash balances should be quite high in developing countries, at least relative to developed countries (see ElGhoul [12]). The short-run adjustment elasticity of 0.86 is double that found for Sudan by Domowitz and Elbadawi [10], and well above the elasticities of 0.52 for the United States found by Gordon [21] and Hendry [29], and the 0.33 and 0.37 reported by Cuthbertson and Hendry [29], respectively, for the United Kindom. The short-run price elasticity is also somewhat higher than the values reported in the aforementioned money demand studies. But the effects of interest rates are low relative to that of the developed countries although in accordance with the general findings of developing countries’ studies. It should be noted though, that in the abovementioned studies, the long-run equilibrium structure was less restrictive than the one used here.

Thus, we have developed a dynamic error correction model based on equation (1) which was designed to satisfy the various evaluation criteria. The model stands up well to the diagnostic tests and is theoretically quite sound. However, in order to judge how ‘good’ the model is it has to be compared with other models used to explain the money demand relationship. It has already been said in the introduction of this paper that models employed to examine the demand for money in the Caribbean have been plagued with serial

---

17 In fact, ElGhoul argues that it should be greater than the unit normally assumed by researchers for developed economies. His thesis is based on precautionary motives for holding money; with a high degree of instability, uncertainty and financial market imperfections, the income elasticity should be very high.
correlation, suggesting that they are misspecified. To illustrate this point further we make a comparison with the study of Bourne [2], which is the seminal work on the demand for money in Jamaica, Bourne postulates the following long-run model:

\[
(1^*) \quad M_t = C_0 + C_1 P_t + C_2 Y_t + C_3 R_t + C_4 \hat{INF}_t
\]

with the following restrictions imposed: \( C_1 = 1 \); \( \hat{INF}_t = \hat{INF}_t \), where the hat(\( \hat{\} \)) represents expectations.\(^{18}\) If we carry out a cointegration exercise on (1*) this model is shown to give non-stationary residuals. This is due primarily to the restriction \( C_1 = 1 \), which renders a DF of -4.249 compared to -5.265 with \( C_1 \neq 1 \). Even if we assumed that the restriction holds, the fact that the foreign interest rate is significant in the ECM of equation (1) would also suggest that Bourne's model is not preferable, despite its parsimony feature.

Bourne's model provides a case for nested hypothesis testing. However, a more interesting and appealing comparison would be with models that are rival to the one presented here. Domowitz and Elbadawi [10] presented such an alternative for the open economy of Sudan which could offer some insights for the Jamaican case, since both economies are quite similar in structural and institutional features. Domowitz and Elbadawi (DE) postulated that:

\[
(6) \quad M = K_0 + K_1 P + K_2 Y + K_3 e + K_4 INF;
\]

where \( e \) is the exchange rate measured as the Jamaican dollar vis-a-vis the US dollar.\(^{19}\) \( K_{ij} \) are elasticities; \( K_0 \) is a constant term and the other variables are defined as before; \( K_4 \) is expected to be negative. The inclusion of the exchange rate captures the direct opportunity cost effect of holding foreign exchange as an alternative to domestic real balances. A change in the return to foreign currency may also have an indirect effect on the transaction component of money demand that operates via the price and inflation variables. So to the extent that foreign currency is considered as an important alternative to domestic money in the wealth portfolio, omission of it may bias a model towards overstating the impact of inflation, in the context of a domestic currency devaluation. Such a variable may be warranted in the Jamaican case, in light of several devaluations over the sample period and recent shortages in foreign exchange. See Worrell [54] for a description of the Jamaican exchange rate system and its recent economic behaviour.

The Granger-Engle two-step method is used to estimate equation (6). The results were as follows:

---

\(^{18}\) Bourne went on from here to specify a dynamic model that included \( M(-1), R(-1), INF(-1) \) and \( Y(-1) \). The result from this model turned out to be inconsistent and, in many respects, implausible, e.g. income elasticities were negative. Notwithstanding this, we concentrate in particular on Bourne's long-run formulation of the demand for money.

\(^{19}\) The source for the exchange rate data is IFS line af. The market rate/par measure is used.
\( M = -2.993 + 1.011Y + 1.246P + 0.125\text{INF} - 0.195e \)

\[ R^2 = 0.997 \quad \text{SER} = 0.072 \quad \text{DW} = 1.929 \quad \text{DF} = -5.705 \]

(ACF) LAG 1-11: 
-0.015 -0.120 -0.218 -0.120 0.098
0.002 -0.221 -0.028 0.438 -0.086 -0.065

\( DM = 0.057 + 0.683\text{DY} + 0.662\text{DP} + 0.447\text{DINF} \)

\[ (0.031) \quad (0.274) \quad (0.245) \quad (0.302) \]

\[-0.530\text{DY}(-1) - 0.964\text{ECMT1}(-1) \]

\[ (0.273) \quad (0.204) \]

\[ R^2 = 0.595 \quad \text{SER} = 0.066 \quad \text{DW} = 1.677 \quad \text{LMN}[X^2(2)] = 0.190 \]

\( HT1[F(1,28)] = 0.019 \quad HT2[X^2(1)] = 1.081 \quad SC1[F(1,23)] = 0.190 \)

\( SC2[X^2(10)] = 9.169 \quad \text{RESET}[F(3,26)] = 1.096 \quad \text{WUT}[F(2,22)] = 2.102 \)

\( \text{CHOW}[F(6,20)] = 1.459 \quad PC1[F(2,24)] = 1.528 \)

The signs and magnitudes of the common variables of equations (4) and (8) are quite similar. Moreover, the DF and the ACF point to \( I(0) \) residuals, despite there being a spike at the ninth lag, supporting a result of cointegrability.\(^{20}\)

The Ljung Box statistic is 11.183. As far as the rank of the cointegrating matrix is concerned, it appears to be of order three with the \( P \), \( M \) and \( Y \) equations all achieving \( I(0) \) residuals (see Table(4)). The \( R^2 \) criterion of Hall [23], however, would suggest continued estimation of the money demand equation, but as before, marginally so.

**TABLE 4. JAMAICA: Cointegrability results of the money demand function.**

<table>
<thead>
<tr>
<th>Inversions</th>
<th>( R^2 )</th>
<th>( DW )</th>
<th>( DF )</th>
</tr>
</thead>
<tbody>
<tr>
<td>( M = -2.99 + 1.25P + 1.00Y + 0.13\text{INF} = 0.20e )</td>
<td>0.9973</td>
<td>1.93</td>
<td>-5.70</td>
</tr>
<tr>
<td>( P = 2.42 + 0.78m - 0.79Y - 0.05\text{INF} + 0.18e )</td>
<td>0.9971</td>
<td>1.89</td>
<td>-5.50</td>
</tr>
<tr>
<td>( Y = 3.02 - 1.14P + 0.91M + 0.01\text{INF} + 0.18e )</td>
<td>0.950</td>
<td>1.71</td>
<td>-5.38</td>
</tr>
<tr>
<td>( \text{INF} = -0.18 + 0.008Y - 0.05P + 0.08M + 0.02e )</td>
<td>0.613</td>
<td>1.33</td>
<td>-3.83*</td>
</tr>
<tr>
<td>( e = -0.54 + 0.23\text{INF} + 1.72Y + 2.49P - 1.66M )</td>
<td>0.91</td>
<td>1.11</td>
<td>-3.25+</td>
</tr>
</tbody>
</table>

*) indicates serial correlation in the \( DF \) formulation.
(+) indicates heteroscedasticity in the \( DF \) formulation.

The parsimonious acceptable \( ECM \) data of equation (7) is shown in equation (8). The diagnostic tests suggest that the disturbance term satisfies the classical

\(^{20}\) \( e \) is \( I(1) \). The \( DF \) value is -5.023.
Furthermore the correlation matrix (Table 5) reveals that the variables are approximately orthogonal, suggesting that multicollinearity is not a serious problem. The coefficients $DY$, $DP$ and the error correction term, ECM(-1), are significant at the conventional level; $DY$ and $DP$ coefficients have the expected sign and are of plausible magnitude. The significance of the ECM term supports the previous conclusion that the variables are cointegrated. $DY(-1)$ is significant at the 10% level and its sign is negative. The exchange rate, although significant in the long-run formulation, is statistically insignificant in the ECM, suggesting that $e$ does not influence the short-run fluctuation of the demand for money. This may reflect the fact that the exchange rate varied very infrequently over the early part of the sample period.

**FIGURE 3. NARROW MONEY: ACTUAL & FITTED VALUES**

![Graph showing Actual and Fitted Values](image)

**FIGURE 4. CUMULATIVE SUM OF SQUARES OF RECURSIVE RESIDUALS**

![Graph showing Cumulative Sum of Squares](image)

The straight lines represent critical bounds at 5% significance level.

---

21 The Wu-Hausman test is calculated using $DY$ and $DP$. The instrumental variables are $Y_{t-1}$, $Y_{t-2}$, $P_{t-1}$, $P_{t-2}$, $e_{1,t}$ and $e_{2,t}$. The plots of the recursive residuals and the fitted and actual are given in figures (3) and (4) respectively.
TABLE 5. ESTIMATED CORRELATION MATRIX OF VARIABLES FOR EQUATION (8)

<table>
<thead>
<tr>
<th></th>
<th>DY</th>
<th>DY(-1)</th>
<th>DP</th>
<th>DINF</th>
<th>ECMT1(-1)</th>
</tr>
</thead>
<tbody>
<tr>
<td>1</td>
<td>1</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>0.457</td>
<td>1</td>
<td>0.551</td>
<td>1</td>
<td></td>
<td></td>
</tr>
<tr>
<td>0.631</td>
<td>-0.107</td>
<td>0.342</td>
<td>1</td>
<td></td>
<td></td>
</tr>
<tr>
<td>0.217</td>
<td>-0.089</td>
<td>0.293</td>
<td>0.432</td>
<td>1</td>
<td></td>
</tr>
</tbody>
</table>

Now we have two tentatively adequate models which satisfy the diagnostic criteria prescribed and, in general, satisfy theoretical requirements. Since these two open economy money demand models are non-nested, statistical methods for testing non-nested hypotheses should be sought. The non-nested tests employed are the F test, calculated by embedding the null and alternative models in a comprehensive model, and the adjusted Cox-type tests, NT and W, which incorporate small sample corrections designed to bring actual significance levels close to the normal values (see Godfrey and Pesaran [20]). Simulation evidence by Godfrey and Pesaran revealed that the NT and W tests have finite sample significance levels which are quite close to the normal value over a wide range of DGP. Moreover, these tests are more likely to lead to the correct decision of accepting the true model and rejecting the false model than the familiar F tests based upon the comprehensive model or regression parameter encompassing approaches. Note, however, that they only test variance encompassing\(^{22}\) and have the drawback that both models can be rejected. Parameter encompassing, theoretically, rectifies this negative aspect of the research and can be tested easily by transforming the more difficult Wald statistic to the classical F test of zero additional parameters in the ‘nested’ model which embeds the maintained and the alternative hypotheses in a linear regression equation. The test statistics are shown below in Table 6.

TABLE 6. NON-NESTED AND ENCOMPASSING TESTS FOR EQUATIONS 5 AND 8.

<table>
<thead>
<tr>
<th>Test Statistic</th>
<th>Distribution</th>
<th>(5) against (8)</th>
<th>Distribution</th>
<th>(8) against (5)</th>
</tr>
</thead>
<tbody>
<tr>
<td>NT-test</td>
<td>N(0,1)</td>
<td>-1.2323</td>
<td>N(0,1)</td>
<td>-1.2750</td>
</tr>
<tr>
<td>W-test</td>
<td>N(0,1)</td>
<td>-1.1125</td>
<td>N(0,1)</td>
<td>-1.1427</td>
</tr>
<tr>
<td>Encompassing F-test</td>
<td>F(3,20)</td>
<td>1.3078</td>
<td>F(4,20)</td>
<td>1.1417</td>
</tr>
</tbody>
</table>

\(^{22}\) This is necessary but not sufficient for encompassing, since parameter encompassing should also be evaluated (see Hendry [28]).
Equation (8) and equation (5) reject each other with NT and W. Thus the negative aspects of the research implied by non-nested tests are manifested. However, the encompassing test presents no solution. According to this test, the null and alternative hypotheses respectively are valid restricted forms of the embedding model which include both of them. Therefore we cannot say that one model encompasses the other. The non-rejection of the models when they may be false may be due to the relatively low power of the F-test when applied to testing non-nested models with a large number of non-overlapping (common) variables (Godfrey and Pesaran [20]). However, the adjusted Cox test does not possess this feature (at least asymptotically).

V. CONCLUSION

This study has attempted to estimate the demand for money function in Jamaica using the recently developed econometric theory of cointegration. It is shown that the dynamic modelling approach provides a model which is a tentatively adequate conditional characterisation of the data generation process. The results of the error correction model show prices, real income, money demand lagged and foreign interest rates to be the main explanatory variables. Inflation does not appear to be a good proxy for the opportunity cost of real assets, at least in the short run. Furthermore our model is shown, through the use of nested procedures, to be better than the earlier formulation of Bourne [2] and in general represents a great improvement on money demand models in the Caribbean which suffer from misspecification. However, non-nested and encompassing tests could not claim preferability for it over the model presented by Domowitz and Eldabawi [10] or vice versa. An interesting research agenda for the future would be to see how these models perform against feed forward models of the demand for money. The latter has the theoretical appeal of being invariant to the so-called Lucas critique, it may not however be realistic in the context of small open economies where it is argued that individuals behave myopically.

BIBLIOGRAPHICAL REFERENCES


---

23 The only common variables are $DP$, $DY$ and the constant.


