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A COINTEGRATION AND ERROR CORRECTION APPROACH TO DEMAND FOR MONEY IN FIJI: 1971-2002*

B. Bhaskara Rao and Rup Singh

University of the South Pacific, Suva (Fiji)

Abstract

Demand for money is an important macroeconomic relationship. Its stability has implications for the choice of monetary policy targets. This paper estimates demand for narrow money in Fiji and evaluates its robustness and stability. It is found that there is a well determined stable demand for money in Fiji, for three decades, from 1971 to 2002 and its dynamics are adequately captured by the cointegration and error-correction models. Income and interest rate elasticities are found to be significant.

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1. INTRODUCTION

Empirical work on the demand for money continues with renewed vigor, for several reasons, in spite of some well established stylized facts about the income and interest rate elasticities. Firstly, demand for money and its stability have important implications for the selection of the monetary policy instruments and for the conduct of monetary policy. According to Poole (1970) interest rate should be selected as the monetary policy instrument when the LM curve is unstable and money supply should be the instrument when IS is unstable. If the choice of the monetary policy instrument is inappropriate then monetary policy will increase the costs of stabilization. Since instability in the demand for money is a major factor contributing to instability in the LM, it is important to test for the stability of the demand for money. Many developed countries have switched to interest rate the as monetary policy instrument when their money demand functions have become unstable following the financial reforms from the second half of the 1980s. However, many developing countries have also abandoned controlling money supply and have been using the rate of interest as monetary policy instrument, even though there is no significant evidence that their demand for money functions have become unstable. Secondly, estimates of the demand for money are useful to understand the limits to non-inflationary seignorage revenue and for the formulation of monetary policy targets. Thirdly, the unit roots and cointegration literature has made significant impact on modeling dynamic economic relationships and especially on the demand for money. Thus, there have been a large number of empirical studies, in both the developed and developing countries, to reestimate demand for money and to investigate, afresh, its stability; see Sriram (1999) for a survey. Fourthly, following Perron's (1989) influential work that the standard unit roots tests lose power if the variables undergo structural changes, lead to a variety of developments in testing for unit roots and estimation of the cointegrating equations.

In comparison to a large number of empirical works on the demand for money for other countries, there are only a handful of empirical studies in Fiji. Furthermore, these existing works are difficult to obtain and seem to have limitations, both in the specification and estimation of the relationship. Therefore, in this paper we review only two more recent empirical works on Fiji with a view to provide a starting point for further work and highlight key issues for further investigation. At the outset, it should be stated that the scope of our paper is not exhaustive because a single paper is not adequate to examine and resolve all the relevant issues.

This paper is organized as follows: In section 2, a brief survey of two recent empirical works on the demand for money in Fiji is given. Our empirical results are discussed in sections 3 to 5. Section 3 presents results of unit root tests. Two alternative methods of estimation of the cointegrating relationship and the corresponding parsimonious short run dynamic adjustment equations are discussed in sections 4 to 5. These are, respectively, estimates based on the General to Specific (GTS) and the Johanson maximum likelihood (JML) approaches. GTS is also known as the LSE-Hendry approach and recently there has been a revival of interest in this approach; see Smith (2000). Finally, our conclusions and limitations are stated in section 6.

2. EMPIRICAL STUDIES IN FIJI

Some earlier works on the demand for money in Fiji like the IMF study in 1982, Luckett (1987) and an unpublished study by Joynson (1997) are hard to obtain. A brief review of these works withtheir limitations are in Katafono (2001). To conserve space these are not reviewed here. More recently there have been two studies. Jayaraman and Ward (2000) have estimated a quarterly model for the demand for broad money and found that it is stable for the period 1979(Q1) to 1996(Q4). Their estimates of the long run income and real interest rate elasticities were 0.987 and +0.022 respectively. Income elasticity was insignificant with a t ratio of 1.33 and the sign of the real interest rate elasticity, measuring the effect of return on quasi money, is positive and significant with a t ratio of 2.05.[1] Jayaraman and Ward argued that the coefficient of the real rate of interest was positive because its positive effect, as the return on quasi-

[1] Inclusion of the real rate of interest in the demand for narrow or broad money is difficult to justify because nominal rates of returns on various liquid assets and their close substitutes are all equally affected by inflation, leaving the relative rates unchanged. Therefore, in several reputed works on the demand for money, including in the textbook discussions, only nominal rates of interest are included; see Friedman (1969), Laidler (1969), and Hendry and Ericsson (1991a and 1991b), Mishkin (2002) etc. The effects of high rates of inflation on liquid asset holdings should be captured by including the expected rate of inflation as a separate variable. It is to be expected that this coefficient will be negative. Inclusion of the real rate of interest implies that the coefficient of the expected rate of inflation is positive. The aforesaid IMF study of 1982 and Jayaraman and Ward (2001) include the real rate of interest. A similar approach was also used by Ahmed (2001) for the demand for money of Bangla Desh. This might be due to the mistaken notion that since the demand for money depends on the real income, it should also depend on the real rate of interest.

money, seems to have dominated its negative effect, as the cost of holding narrow money. In her review Katafono (2001) has further pointed out that the finding that the demand for money in Fiji is stable by Jayaraman and Ward, with generated quarterly GDP data, is of little use for policy because it is not possible to forecast the demand for money since quarterly estimates of GDP are not available in Fiji. Furthermore it is hard to accept their conclusion that the monetary authorities in Fiji should use broad money as the monetary policy instrument because their estimated income elasticity is insignificant.

Given such limitations in the earlier works, the more recent study by Katafono (2001) stands out as a significant contribution and in our view a good starting point for further work in Fiji and other Pacific Island Countries (PICs). Therefore, in the rest of this section, we shall review her work in some detail. Katafono has applied time series econometric techniques and her systematic approach is refreshing. She has estimated the demand functions for narrow, quasi and broad monies in Fiji for the period 1975 to 1999, utilizing the existing annual data. However, the demand for narrow money, i.e., M1 as it is commonly known, has received relatively more attention in her work. Since our objective is to estimate and analyze demand for M1, we only review this part of her work.

Katafono has used a standard specification of the demand for money i.e., M1 (henceforth M), in a semi logarithmic form:

$$ln\Big(\frac{M_t}{P_t}\Big) = \alpha_0 + \alpha_1 lnY_t + \alpha_2 SVR_t + \alpha_3 TBR + \alpha_4 lnREER_t + \epsilon_t \qquad (1)$$

where M is nominal money, P is price level (CPI), SVR is the nominal rate of interest on saving deposits, TBR is the nominal treasury bill rate, REER is the real effective exchange rate and ϵ is an *iid* error term.[2]

After the unit root tests showed that these are all I(1) variables, she has conducted cointegration tests on these variables with the Johansen maximum likelihood method (JML) and found that there is one cointegrating vector and interpreted it as the demand for money after conducting the usual causality tests. However, these tests did not conclusively establish that money does not Granger cause the two interest rates. In spite of this, as is common in many empirical works on the demand for money, she has interpreted the cointegrating vector as the demand for money because no other sensible alternative is plausible. The long run equilibrium money demand function implied by the JML approach is:[3]

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[2] Both TBR and SVR should be treated as the opportunity cost of holding narrow money although their coefficients are unlikely to be determined well due to multicollinearity between these two rates of return.

[3] We are grateful to Katafono for pointing out that, in an earlier version of our

$$ln\left(\frac{M_t}{P_t}\right) = -\,2.964 + 0.610\ lnY_t - 0.190\ SVR_t + 0.104\ TBR - 0.048\ lnREER_t \tag{2}$$

Instead of estimating the short run dynamic adjustment relationship with the lagged residuals of equation (2), where the parameters in the cointegrating equation are estimated efficiently with a systems method, she has estimated, afresh with OLS a variant of this equation, based on the well known General to Specific approach (GTS) of Hendry.[4]

In Katafono's two estimates there are some minor differences in the estimated long run coefficients with GTS and JML. It would have been valuable if she has estimated a parsimonious dynamic adjustment equation with the lagged residuals from her cointegrating equation for comparisons. She did not use this option and estimated, a fresh, her GTS version with OLS. The implied long run relationship of her GTS estimates is:[5]

$$ln\left(\frac{M_t}{P_t}\right) = 0.511\ lnY_t - 0.104\ SVR_t + 0.004\ TBR - 0.150\ lnREER_t \tag{3}$$

It can be seen that there are only small differences in the income and interest rate elasticities obtained with JML in equation (2) and GTS in equation (3) and their coefficients are correctly signed. The coefficients of TBR are contrary to expectation although the coefficients of REER are correctly signed. Katafono repeatedly states that the sign of the coefficients of SVR should be positive. This is contrary to what she has found and also with the usual expectation that, in the demand for narrow money, the rate of interest on time and saving deposits as the price of holding money (M1) should in fact be negative. However, the coefficients of TBR are positive in both equations and that may be due to some collinearity between these two rates of interest. Perhaps Katafono should have restimated both her equations by deleting TBR from the specifications. The major problem with her two estimates is that they imply implausibly low

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paper, we did not correctly report the estimated coefficients of her equation based on JML. However, she acknowledged our corrections to her GTS estimates, reported below in equation (3).

[4] See Charemza and Deadman (1997) and Smith (2000) for an exposition of the GTS approach. Katafono might have opted for this single equation GTS because the weak exogenity assumption for the two interest rates is rejected.

[5] There seem to be some typographical errors in the estimates shown in column 2 of Table-4 of Katafono. Therefore, we have adjusted these estimates. See also footnote 3.

income elasticities of about 0.5, contrary to her claim that it is close to unity; see Table 14 in Jayaraman and Ward (2000) for a useful summary of the income elasticities of some developing countries which range from 1.85 for Indonesia to near unity for Fiji. Moreover, her final estimate of the demand for money is also found to be temporally unstable. In conclusion it may be said that in spite of some limitations, Katafono's work provides useful insights and a good starting point for further work.

3. UNIT ROOT TESTS

We first start with the tests for stationarity of the three variables, (M/P), Y and R, in the function for demand for real narrow money balance (M/P):

$$ln\left(\frac{M_t}{P_t}\right) = \alpha_0 + \alpha_1 ln\ Y_t + \alpha_2 R_t + \epsilon_t \tag{4}$$

where M is narrow money consisting of currency in circulation and demand deposits, P is the GDP deflator, Y is the real GDP measured at factor cost and R is the nominal 1-3 years weighted average interest rate on time deposits and ϵ is an *iid* error term. Our sample period extends from 1971 to 2002. Definitions of the variables and sources of data are given in Appendix.[6]

A preliminary estimate of equation (4) using the simple OLS procedure and partial adjustment mechanism gave promising results. These are not reported here to conserve space and also because our unit root tests below show that the three variables in equation (4) are non-stationary in their levels but stationary in their first differences. Therefore, OLS estimates with the levels of these variables give misleading estimates of standard errors and other summary statistics. The unit roots test results for the variables in equation (4) are given in Table-1.

Conventional unit root test statistics based on ADF and PP do not reject the unit root null for the levels of the variables at the conventional 5% or 10% levels. Two other test statistics are used. Pantula *et.al* (1994) developed the weighted symmetric ADF statistic or ADF(WS), which dominates in terms of power over all other tests. This is available in TSP.

[6] In the earlier studies on the demand for money, notably Katafono (2001), the real effective exchange rate was introduced as an explanatory variable without an adequate explanation of whether holding foreign exchange balances, as a substitute for domestic money, is a realistic option in Fiji. If that were a possibility, in addition to the real effective exchange rate, there should be a return variable, e.g. a weighted average of some deposit rates in the trading partner countries. We have ignored, however, this variable because we consider that foreign exchange holdings is not a realistic option in Fiji.

Table 1

Tests for Unit Roots:
Levels and First Differences of Variables
with Intercepts and Linear Trends.

Variable	m	ADF	ADF(WS)	PP	ERS
$ln(M/P)$	$[2,2,2,0]$	-2.26 (0.44)	-0.73 (0.99)	-6.38 (0.71)	15.84 (5.72)
$\Delta ln(M/P)$	$[1,1,1,0]$	-5.17 (0.00)	-5.42 (0.00)	-39.20 (0.00)	1.87 (2.97)
$ln\ Y$	$[2,3,3,1]$	-1.44 (0.81)	-1.87 (0.73)	-16.80 (0.13)	21.28 (5.72)
$\Delta ln\ Y$	$[2,2,2,0]$	-4.30 (0.00)	-2.4 (0.06)*	-37.76 (0.00)	2.63 (2.97)
R	$[2,5,5,2]$	-1.79 (0.61)	-0.90 (0.98)	-2.26 (0.96)	69.34 (5.72)
ΔR	$[1,1,1,0]$	-4.49 (0.00)	-4.75 (0.00)	-32.15 (0.00)	1.56 (2.97)

Notes:

ADF is the standard augmented Dicky-Fuller F test, ADF(WS) is the weighted symmetric ADF test, PP is the Phillips-Perron test and ERS is the Elliott-Rothenberg-Stock test. ADF(WS) seems to dominate other tests in terms of power; see Pantula et.al (1994).

m is the lag length of the first differences of the variable included. For example $[1,1,1,1]$, means that one lagged first difference is found to be adequate in the four test statistics, respectively.

The sample periods chosen for the test are 1972/2002 for the levels and 1973/2002 for the first differences of the variables.

p values are given below the test statistics in parentheses, except for the ERS. For the ERS, the 5% critical values are shown in paranthesis. In E-views, the null hypothesis of unit roots is rejected if the computed ERS test statistic is below the critical value.

A time trend is included in the levels but not in the first differences of the variables. TSP 4.5, Microfit 4.1 and E-views 5.0 are used to estimate the test statistics.

ADF(WS) also shows that the unit root null cannot be rejected for the levels of the variables. Similarly the computed Elliott-Rothenberg-Stock test statistics (ERS) are more than the 5% critical values, implying that all the levels of the variables are non-stationary. However, the p values for the first difference of these variables are all significant at the 5% level and reject the unit root null. The computed test statistics for ERS are also below the 5% critical values. Therefore, these variables are $I(1)$ in levels and $I(0)$ in their first differences.

4. EMPIRICAL ESTIMATES: GENERAL TO SPECIFIC APPROACH

If all the variables are found to be I(1), three different methods can be used to find if they are cointegrated. These are the Engel-Granger (EG) two step procedure, the General to Specific (GTS) approach and the Johansen maximum likelihood method (JML). Maddala and Kim (1998) review these approaches and note that among such alternatives, the LSE-Hendry GTS approach is popular in empirical work because it can be easily implemented. Therefore, in this section we shall first use GTS and in the next section the JML approach.

In GTS, first, a very general dynamic lag structure between the dependent and explanatory variables – consisting of their lagged levels and first differences – is estimated with OLS. In the second stage, this overtly general specification is reduced into a parsimonious dynamic adjustment equation, using the variable deletion tests by ensuring that the overall summary statistics do not become significant and reject the null that the residuals satisfy the underlying classical assumptions.[7]

Before we use this technique it is necessary to understand a transformation necessary to give an error correction model (ECM) interpretation to the estimated equation. The basic equilibrium specification of the demand for money, such as:

$$ln\left(\frac{M_t}{P_t}\right) = \alpha_0 + \alpha_1 lnY_t + \alpha_2 R + \epsilon_t \tag{5}$$

can be equivalently written as:[8]

$$\begin{aligned}\Delta ln\left(\frac{M_t}{P_t}\right) =& \beta_0 + \beta_1 ln\left(\frac{M_{t-1}}{P_{t-1}}\right) + \beta_2 lnY_{t-1} \\ &+ \beta_3 R_{t-1} + \beta_4 \Delta ln\left(\frac{M_{t-1}}{P_{t-1}}\right) + \xi_t \end{aligned} \tag{6}$$

7 A good exposition of GTS can be found in Charemza and Deadman (1997). The famous Davidson, Hendry, Srba and Yeo (1978), DHSY for short, work on the consumption function for the UK is now a classic paper on GTS. Subsequently, Hendry and Ericsson (1991a and b) have used this approach to reestimate and test the money demand function for the USA, of Friedman and Schwartz (1982).

8 This formulation is based on Bejerjee, Dolado, Hendry and Smith (1986). For a simpler exposition of this transformation see Cuthbertson (1995).

Although these equation seem simple, they are computationally demanding because the general dynamic specification of equation (6) will include many more lagged values of the relevant variables. Furthermore, there are no clear cut guidelines on how to reduce the long lag structure to arrive at a manageable parsimonious final equation. The general dynamic version of equation (6) can be specified as:

$$\begin{aligned}\Delta ln\left(\frac{M_t}{P_t}\right) =&\beta_0 + \beta_1 ln\left(\frac{M_{t-1}}{P_{t-1}}\right) + \beta_2 lnY_{t-1} + \beta_3 R_{t-1}\\ &+ \sum_{i=0}^{n}\lambda_i \Delta lnY_{t-i} + \sum_{i=0}^{m}\gamma_i \Delta lnR_{t-i}\\ &+ \sum_{i=1}^{j}\tau_i \Delta ln\left(\frac{M_{t-i}}{P_{t-i}}\right) + \xi_t \qquad (7)\end{aligned}$$

It can be seen that this specification retains the error correction part, given by the lagged levels of the variables, and the equilibrium long run coefficients are given by $\beta_0,\ \beta_1,\ \beta_2$ and β_3. If the three I(1) level variables are cointegrated, since their first differences are stationary, the error term ξ will be I(0) and satisfies the standard classical assumptions. Therefore, OLS can be used to estimate equation (7).

However, before we report our estimates, we have added some dummy variables to equation (7). First, a coup dummy variable (COUP) which is 1 since 1988 and zero in all other periods is expected to capture the political uncertainty effect on the demand for money. It is reasonable to expect that its coefficient would be positive because it is likely to increase holdings of precautionary balances. Second, there have been two devaluations in 1987 and 1998. Devaluations cause an anticipated increase in the prices of the imported goods, although there would be some lag between devaluations and the increase in the prices of imported goods. This is the well known exchange pass-through effect. The effects of devaluations, therefore, would be immediate but transitory. Immediately after devaluations, there would be a sudden increase in the purchase of imported goods, causing a shift from holding money to holding real goods. Therefore, the coefficient of this dummy variable (DEV) is expected to be negative in the demand for money function. Finally, the collapse of the National Bank of Fiji in 1996 might have caused loss of confidence and a shift away from bank money in particular. These confidence loss effects seem to have persisted for a while. Therefore, our NBF dummy variable is 1 from 1996 to 1998. To gain a degree of freedom, we have combined the negative effects of the two devaluations and the collapse of the NBF into a single dummy variable DEVNBF.

In Table-2 a few parsimonious versions of equation (7) are reported.

Table-2
GTS Short-Run Adjustment equations

	7a	7b	7c	7d	7e	7f
Intercept	-4.455 (-0.73)	-3.966 (-0.53)	-1.532 (-1.04)	-2.147 (-5.87)*	-2.139 (-5.98)*	-2.047 (-5.34)*
Trend	-0.009 (-0.51)	-0.008 (-0.36)		-0.002 (-0.47)		
$ln\ (M/P)_{t-1}$	-1.151 (-5.07)*	-1.205 (-5.10)*	-1.189 (-5.13)*	-1.199 (-5.31)*	-1.169 (-5.51)*	-1.109 (-5.21)*
$ln\ (Y/P)_{t-1}$	1.475 (1.68)	1.451 (1.38)	1.107 (4.22)*	1.199 (5.31)*	1.169 (5.51)*	1.109 (5.21)*
R_{t-1}	-0.034 (-2.48)*	-0.037 (-2.75)*	-0.036 (-2.79)*	-0.037 (-2.74)*	-0.034 (-2.93)*	-0.031 (-2.53)*
$\Delta ln\ (Y/P)_t$	1.922 (3.03)*	1.785 (2.65)*	1.599 (4.27)*	1.646 (3.78)*	1.646 (3.86)*	1.742 (4.71)*
$\Delta ln\ (Y/P)_{t-4}$	0.816 (2.03)**	0.838 (1.63)*	0.809 (1.68)	0.813 (2.12)*	0.832 (2.23)*	0.802 (1.69)*
ΔR_t	-0.052 (-2.33)*	- 0.055 (-2.70)*	-0.051 (-2.91)*	-0.053 (-2.54)*	-0.049 (-2.64)*	-0.045 (-2.68)*
COUP	0.314 (2.61)*	0.322 (2.45)*	0.280 (3.79)*	0.296 (3.61)*	0.265 (5.41)*	0.247 (4.07)*
DEVNBF		-0.030 (-2.27)	-0.031 (-2.36)*	-0.031 (-1.08)*	-0.031 (-1.12)	
$\overline{R}^2$	0.644	0.645	0.661	0.663	0.677	0.673
SEE	0.087	0.087	0.085	0.085	0.083	0.083
χ^2_{sc}	0.713 (0.40)	0.004 (0.95)	0.111 (0.74)	0.060 (0.81)	0.033 (0.86)	0.265 (0.61)
χ^2_{ff}	4.578* (0.03)	4.329 (0.04)	3.553 (0.06)	3.55 (0.06)	3.33 (0.07)	3.461 (0.06)
χ^2_{n}	0.040 (0.98)	0.088 0.96)	0.087 (0.96)	0.107 (0.95)	0.079 (0.96)	0.178 (0.92)
χ^2_{hs}	3.376 (0.07)	2.827 (0.09)	3.955 (0.05)	3.658 (0.06)	3.741 (0.05)	4.408 (0.04)

t-ratios are in paranthese below the coefficients. For the χ^2 test statistics p-values are in the parantheses. $*$ and $**$ signify 5% and 10% significance levels respectively. In equations where χ^2_{hs} is significant, we have used the Newy-West adjusted standard errors.

In equation 7a, in the second column, all the summary statistics are satisfactory, except that the functional form misspecification χ^2_{ff} (RESET) test is significant at the 5% level but not at 1% level. The t-ratios in the parentheses below the coefficients indicate that the estimated income elasticity is insignificant even at the 10% level. Its p-value (not given in the table), however, is 0.11 implying that it is significant at a slightly higher level. The coefficient of time trend is also insignificant. It is noteworthy that the coefficients of the rate of interest and the coup dummy variable have the expected negative and positive signs, respectively, and are significant. The implied income elasticity, although insignificant, seems to be on a slightly higher side at 1.28. However, this is not unusual for the developing countries; see Table 14 in Jayaraman and Ward (2000) for the estimates of the income elasticities of some developing countries.

When this equation was tested for temporal stability with TIMVAR tests, the CUSUM test indicated instable from 1998 onwards, but the CUSUM SQUARES test showed that it is stable. To improve the summary statistics, we have added a devaluation dummy, for devaluations in 1987 and 1998. This did not improve the results and its coefficient was insignificant. We have then also added a dummy variable for the collapse of the National Bank of Fiji 1996 but this did not improve the results. However, when these two dummies are combined as DEVNBF, there was some improvement in the summary statistics, and estimate of the income elasticity decreased marginally to 1.20 but is significant only at 18.5% level. These results are given in equation (7b) in Table-2. The CUSUM test showed considerable improvement but indicated that there was still some instability in the demand for money since 1998. The CUSUM SQUARES test, however, did not show any temporal instability.

It may be noted that the trend variable remained highly insignificant in both equations. Although it is essential to include a trend variable in the VAR models, plots of real money, real output and the rate of interest show that these variables are not strongly tended in Fiji. Therefore, we tested for the constraint that the coefficient of the trend variable is zero. The computed $\chi^2(1)$ test statistics is 0.164 and significant only at 69%. Therefore, equation (7c) in Table-2 is estimated without the trend. It can be seen that all the summary statistics showed improvements except χ^2_{hs} for heteroscedasticity, which is now significant at 5% but not at 1% level. Three interesting changes are noteworthy. First, the estimate of income elasticity is almost unity. Second, the functional form misspecification χ^2_{ff} statistic is insignificant at 5% level. Third, the Newy-West adjusted standard errors indicate that the devaluation and NBF dummy is significant and has the expected negative sign.

Encouraged by this result we tested for the constraint that the income elasticity of demand for money is unity with and without the trend variable. The computed $\chi^2(1)$ test statistics for this constraint are 0.080 and 0.323 and significant at 78% and 57% respectively. Therefore, equations (7d) and (7e), with and without trend, are estimated with the constraint of unit income elasticity. It can be seen that both these equations are well determined. None of the χ^2 summary statistics are significant at the 5% level and all other coefficients, except those of the dummy variable, are also significant at the 5% level. When these equations are subjected to the TIMVAR tests both the CUSUM and CUSUM SQARES tests indicated no temporal instability.[9]

Finally, since the devaluations dummy is not significant, we deleted it and reestimated equations (7d) and (7e). While this did not make any difference to the estimates of these two equations, the CUSUM test for the equation with trend showed instability from 1998. However, the CUSUM SQUARES TEST showed no instability. In the equation without trend, given as equation (7f) in Table-2, neither stability test showed any instability. It is hard to determine which of these six equations is the best since they all have similar summary statistics and close standard errors of estimates. The SEEs of our equations are similar to those reported by Katafono. The constrained equation (7e) has the lowest SEE of 0.081. However, we prefer equation (7d) because of the presence of the trend variable. The actual and predicted values of the change in the logarithm of real money are plotted in Figure-1 below. It can be seen that the fit is fairly good except for 1978, 1983, 1989 to 1995.[10]

5. COINTEGRATION AND ECM ESTIMATES: JOHANSEN METHOD

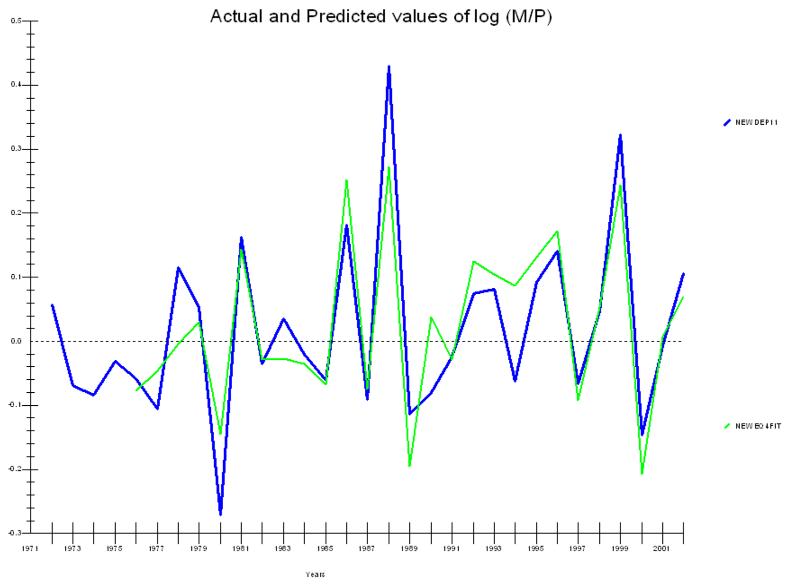
The Johansen (1988) ML procedure (JML) in Microfit is used for testing the existence of the cointegrating relationships in equation (5). We first tested for the optimum lag length of the VAR with a 4^{th} order model, by using the unrestricted VAR model option in Microfit. The I(0) variable selected are intercept, time trend and the two dummy variables used earlier COUP and DEVNBF. The Akaike Information Criterion (AIC) reached a maximum of 42.65 for VAR(2) but the Schwarz Criterion (SBC) reached a maximum of 26.75 for VAR(1). Since our sample size is small, we have selected VAR(1). We postpone the Granger causality tests until we test for the number of the cointegrating vectors. JML estimates implied that the null of no cointegration can be rejected for VAR(1) but not for VAR(2). In VAR(1) the null that the number of cointegrating vectors is zero ($r = 0$),

[9] To conserve space the TIMVAR plots are not reported but they can be obtained from the authors.

[10] A regression between the actual and fitted values showed that the intercept is zero ($-0.91E^{-7}$) and the slope is unity. However, $\overline{R}^2 = 0.755$ and SEE=0.072.

FIGURE 1

ACTUAL AND PREDICTED VALUES FORM EQUATION 7d



is rejected by the trace test statistic only at the 10% level. The computed value is 28.98 and the 10% critical value is 28.78. The null that $r = 1$ is accepted by the eigenvalue and trace test statistics at the 95% level. The single cointegrating vector, normalized on money, obtained with JML is:

$$ln\left(\frac{M_t}{P_t}\right) = 1.133 ln\ Y_t - 0.037 R_t \tag{8}$$

We have conducted weak and strong exogenity tests for the null that money does not Granger cause income and the rate of interest. The computed $\chi^2(2)$ test statistic for the weak exogenity test, with p value in the parenthesis, is 6.04 (0.049). The corresponding strong exogenity test statistics $\chi^2(4)$ is 11.59 (0.021). In both cases the null can be accepted only at the 1% but not at the 5% level. Subject to these limitations, it is reasonable to interpret this single cointegrating vector as the demand for money. Therefore, in equation (8) the cointegrating vector is normalized on real money. The two crucial coefficients of income and rate of interest have the expected signs and magnitudes. The estimated income elasticity of demand for money is almost unity at 1.133, in comparison to an estimate of about 0.7 estimated by Katafono. The implied interest elasticity, at the mean interest rate of 6.97 is -0.286 is also plausible. These elasticities are comparable to similar recent estimates for India by Pradhan and Subramanian (2003) and Hafer and Kutan (2003) for Philippines.

In developing the ECM model we adopted the GTS approach in the second stage. The second stage equation can be estimated with OLS using the lagged residuals from the cointegrating vector of JML. Estimation of ECM with OLS does not lead to biased estimates because the second stage equation puts no restrictions on the first stage cointegrating vectors.[11]

In all the following equations t values are given below the coefficients in parentheses and p values are given in the parentheses for the χ^2 summary statistics. 5% and 10% significance is indicated with * and ** respectively.

$$\begin{aligned}
\Delta ln\left(\frac{M_t}{P_t}\right) = & -2.952 - 0.002\ T - 1.079\ ECM_{t-1} + 1.136 \Delta lnY_t - 1.246 \Delta lnY_{t-1} \\
& (-6.38)^{*}(-0.47) \quad (-6.22)^{*} \qquad (3.07)^{*} \qquad (-2.86)^{*} \\
& -0.827 \Delta lnY_{t-2} - 0.826 \Delta lnY_{t-4} - 0.039 \Delta R_t + 0.035 \Delta R_{t-1} \\
& (-2.18)^{*} \qquad (2.84)^{*} \qquad (2.62)^{*} \qquad (3.08)^{*} \\
& + 0.269 COUP - 0.119 DEVNBF \\
& (4.36)^{*} \qquad (1.76)^{**} \qquad\qquad (9)
\end{aligned}$$

$$\overline{R}^2 = 0.825, SEE = 0.061 \quad Period: 1976 - 2002$$
$$\chi^2_{sc1} = 0.411\ (0.52), \chi^2_{ff} = 5.68^{*}\ (0.02)$$
$$\chi^2_{n} = 0.684\ (0.71), \chi^2_{hs} = 0.780\ (0.38)$$

The summary statistics of this equation are good and a noteworthy feature of this equation is that it has a lower SEE of about 0.06 compared to 0.08 in all earlier estimates, including the estimates by Katafono. However, it may be noted that the functional form misspecification χ^2_{ff} test is significant at 5% but not at 1% level. This is not unusual for dynamic equations because it is hard to claim that the complex nature of dynamic adjustments, with limited data, can be adequately captured with linear specifications. All the coefficients are significant except that of the time trend. The combined devaluation and NBF dummy is significant at 10% level. When we tested separately for the significance of the two devaluations in 1988 and 1997 as well as the failure of the National Bank of Fiji, the second devaluation seems to have had a larger impact. Therefore, the above equation is reestimated with a dummy variable only for the second devaluation, its coefficient is significant at the 10% level. The functional form misspecification test statistic deteriorated somewhat but still insignificant at the 1% level. These estimates are given below.

[11] One of the referees has suggested that it is desirable to estimate the ECM with the systems method. In Microfit the second stage equations are actually estimated with OLS, using the lagged ECM part from the first stage. However, the order of the actual dynamic equations is limited to the chosen first order. In our view this procedure unnecessarily restricts the order of the dynamic second stage equation.

$$
\begin{aligned}
\Delta ln\Big(\frac{M_t}{P_t}\Big) =& -2.952 - 0.002\ T - 1.079\ ECM_{t-1} + 1.136\Delta lnY_t - 1.125\Delta lnY_{t-1} \\
& (-6.38)^{*}\ (-0.475)\quad (-6.22)^{*}\qquad (3.07)^{*}\qquad (-2.86)^{*} \\
& - 0.827\Delta lnY_{t-2} + 0.827\Delta lnY_{t-4} - 0.038\Delta R_t + 0.035\Delta R_{t-1} \\
& (-2.18)^{*}\qquad (2.84)^{*}\qquad (-2.62)^{*}\qquad (3.08)^{*} \\
& + 0.269COUP - 0.119DEV2 \\
& (4.36)^{*}\qquad (-1.75)^{**}
\end{aligned}
\tag{10}
$$

$$\overline{R}^2 = 0.825, SEE = 0.06 \quad Period : 1976 - 2002$$

$$\chi^2_{sc1} = 0.411\ (0.52), \chi^2_{ff} = 5.68^{*}\ (0.02)$$

$$\chi^2_{n} = 0.684\ (0.71), \chi^2_{hs} = 0.780\ (0.38)$$

It may be noted from these estimates in equations (9) and (10) that it is possible to reduce further the number of estimated coefficients to increase the degrees of freedom. The positive coefficient of ΔlnY_t is close in value to the absolute values of the coefficient of ECM_{t-1} and ΔlnY_{t-1}. Furthermore, the coefficients of ΔR_t and ΔR_{t-1} are close and opposite in sign. Similarly the coefficients of ΔlnY_{t-2} and ΔlnY_{t-4} are close but opposite in sign. When these four restrictions are tested, the computed $\chi^2(4)$ test statistic with p value in the parenthesis is 0.371 (0.985) and insignificant. Therefore, this constraints could not be rejected. The following ultra parsimonious equation is based on these restrictions:

$$
\begin{aligned}
\Delta ln\Big(\frac{M_t}{P_t}\Big) =& -3.047 - 0.002\ T - 1.114\ ECM_{t-1} + 1.114(\Delta lnY_t - \Delta lnY_{t-1}) \\
& (-10.94)^{*}\ (-0.706)^{*}\quad (-11.05)^{*}\qquad (11.05)^{*} \\
& - 0.820(\Delta lnY_{t-2} - \Delta lnY_{t-4}) - 0.039(\Delta R_t - \Delta R_{t-1}) \\
& (-3.59)^{*}\qquad\qquad (6.43)^{*} \\
& + 0.279COUP - 0.114DEV2 \\
& (5.93)^{*}\qquad (1.93)^{**}
\end{aligned}
\tag{11}
$$

$$\overline{R}^2 = 0.857, SEE = 0.055 \quad Period : 1976 - 2002$$

$$\chi^2_{sc1} = 0.169\ (0.68), \chi^2_{ff} = 3.94^{*}\ (0.05)$$

$$\chi^2_{n} = 1.053\ (0.59), \chi^2_{hs} = 0.642\ (0.42)$$

The summary statistics of this equation are impressive and the estimated coefficients are similar to those in the previous two equations. There is a marginal reduction in the SEE from 0.06 to 0.055. When this equation was tested for temporal stability neither the CUSUM nor CUSUM SQUARES test showed any instability. The plots from these two tests are given below in Figure-2 and Figure-3. The predicted and actual values from equation (11) are plotted in Figure-4.

FIGURE 2

CUSUM TEST FOR EQUATION 11

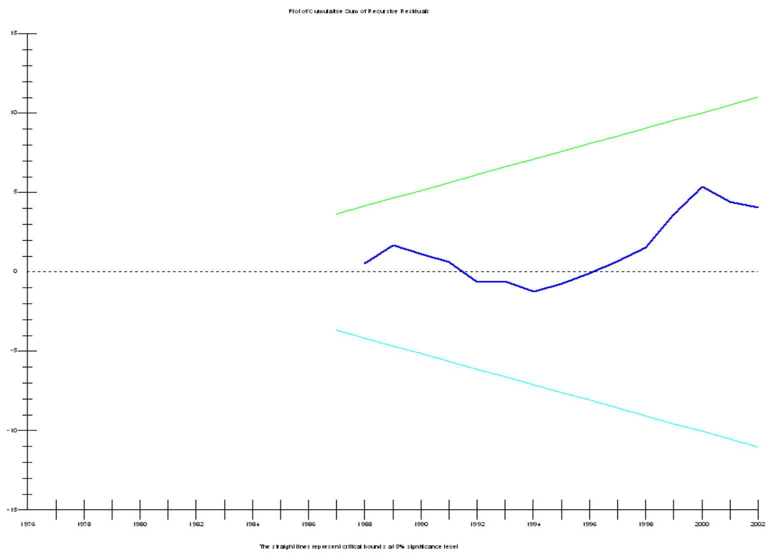


FIGURE 3

CUSUM SQUARES TEST FOR EQUATION 11

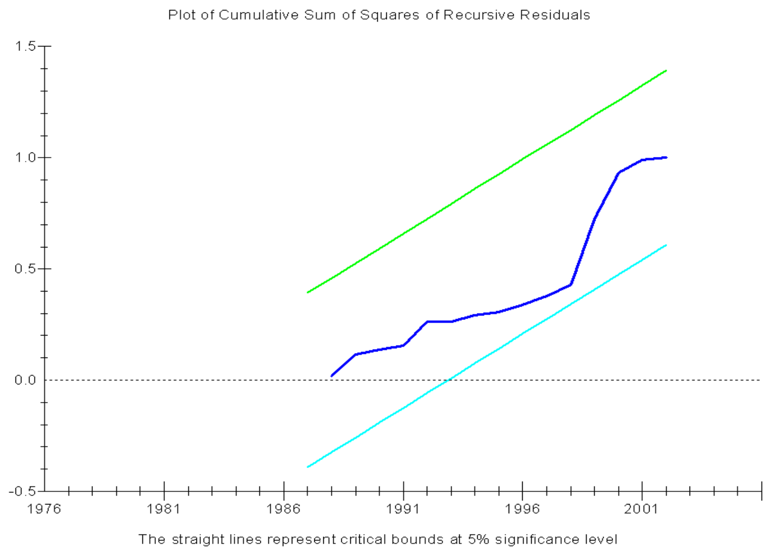
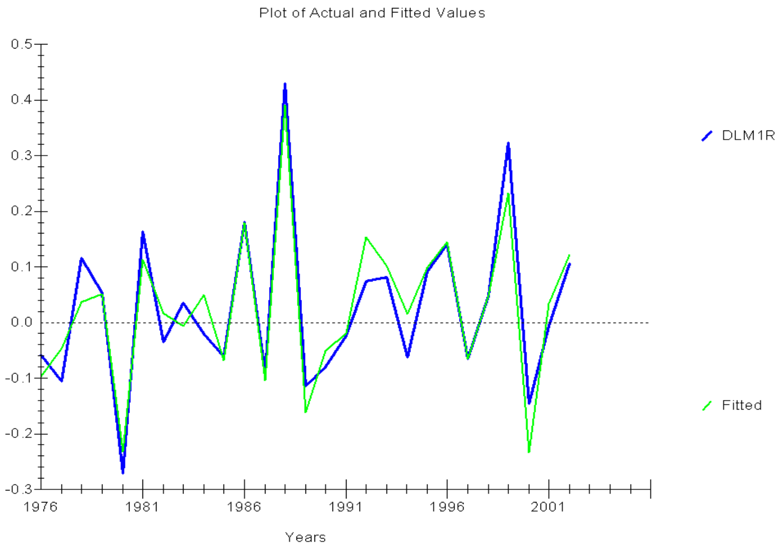


FIGURE 4

ACTUAL AND FITTED VALUES FROM EQUATION 11



6. CONCLUSIONS

In this paper, we have surveyed earlier works on the demand for money in Fiji. It is noted that Katafono's work has many merits and is a good starting point for further work. However, while Katafono's work is relatively free from some weaknesses in the other earlier works, it is also found to be in need of further improvements. Therefore, we have used two alternative methods of estimation of the demand for money in Fiji, using the time series econometrics methods. The GTS and JML method have yielded similar cointegrating coefficients although their dynamic adjustment lags are somewhat different. The estimated income and interest rate elasticities are found be well determined and their signs and magnitudes are consistent with *prior* expectations.

Our first major finding is that in Fiji the income elasticity of the demand for narrow money (M1) is unity and the interest rate elasticity is negative and about -0.35. Our second major finding is that the demand for money in Fiji is temporally stable. Therefore, our work raises doubts on the appropriateness of Reserve Bank of Fiji's monetary policy of targeting the rate of interest, instead of the stock of the real narrow money balances.

Some caveats about our findings should be also noted. First, several test statistics we have used are appropriate only for large samples. Therefore, in any further work it is important to make some adjustments to minimize our finite sample biases. Such adjustments need considerable computational effort and they falls outside the scope of our present paper. Second, we have ignored structural breaks and their implications for the unit root tests and estimation of the cointegrating relationships. It may be noted that the main contribution of the literature on structural breaks is to improve the power of the unit root tests. If there is support for structural breaks then there are two ways of proceeding further with estimating cointegrating vectors. First, if the unit root tests, with structural breaks, show that there are no unit roots in the variables, then the relationships can be estimated with the classical methods with appropriate shift dummies; see Rao (1993a) and (1993b) for an early application of this approach. Second, if in fact there are unit roots in the variables, cointegrating relationships are generally estimated for various sub-periods during which there are no structural breaks; see Choi and Jung (2003) for a recent application of this procedure. The Choi-Jung procedure requires a large number of observations, in each sub-sample, to get any meaningful cointegrating relationships and therefore not useful for developing countries with limited number of annual data.[12]

Third, we have ignored the demand for broad money and its stability. Given these limitations, our findings should be treated only as the maintained hypotheses until they are refuted by other works. Consequently, before we recommend our findings for policy formulation in Fiji, we emphasize, in no uncertain terms, the need for further work based on more refined techniques and better insights into the theory of the demand money. We only hope that our work and methodology will be useful, together with Katafono's work, as starting points for further work in Fiji and other PICs.

---

[12] While the theoretical developments in structural changes are valuable, it may be noted that Maddala and Kim take a cautious view about their practical use with the following observation:

> "There is a lot of work on testing with unknown switch points. In practice, there is a lot of prior information and there is no reason why we should not use it. For instance, suppose there is a drastic policy change or some major event (for example, oil price shock) that occured at time t_0. *It does not make sense* to ask the question of whether there was a structural change around that period. *It is not very meaningful* to search for a break over the entire sample period ignoring this prior information." Maddala and Kim (1998, p.398), our italics.

These observations imply that perhaps testing for unit roots with *a priori* known dates, e.g. Perron (1989), is more meaningful than the more recent approaches based on endogenous switching points. Needless to say this is a philosophical issue and therefore there are likely to be many views.

Data Appendix

P = GDP deflator. Is a ratio of nominal to real GDP in 1995 prices. Source: International Financial Statistics (2003 CD-ROM) and the Reserve Bank of Fiji Quarterly Review (various years).

Y = GDP at factor cost in 1995 prices. Source: Reserve Bank of Fiji Quarterly Review (various years) and the International Financial Statistics (2003 CD-ROM).

R = Nominal interest rate. Is the simple average of 1-3 years savings deposit rate. Source: Reserve Bank of Fiji Quarterly Review (various years).

M1 = Narrow money balance. This includes currency in circulation, demand deposits and bills payable. Source: Reserve Bank of Fiji Quarterly Review (various years) and the International Financial Statistics (2003 CD-ROM).

COUP = dummy variable for the two political coups in Fiji. Data constructed as 1 since the first coup in 1987 up till 2002 and 0 in all other periods.

NBF = dummy variable for the collapse of the National Bank of Fiji. Data constructed as 1 for 1996 to 1998 and 0 for all other periods.

DEV = dummy variable for the two devaluations of the domestic currency. Data constructed as 1 for 1987 and 1998 and 0 for all other periods.

DEVNBF = dummy variable for the devaluations and the National Bank of Fiji crisis. Data constructed by adding the individual dummy variables.

Notes:

1. All variables, except the rate of interest and dummies, are deflated with the GDP deflator and are converted to natural logs.
2. Data are available for replication on request.

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