A cointegration and error correction approach to demand for money in Fiji, 1971–2002

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The stability of the demand for money has implications for the choice of monetary policy targets. This paper estimates the demand for narrow money in Fiji and evaluates its stability. It is found that there is a well-determined stable demand for money for the period 1971 to 2002 and its dynamics are adequately captured by cointegration and error-correction models. Income and interest rate elasticities are found to be significant.

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Empirical work on the demand for money continues with renewed vigour for several reasons, in spite of some well-established stylised facts about the income and interest rate elasticities. First, the demand for money and its stability have important implications for the selection of monetary policy instruments and for the conduct of monetary policy. According to Poole (1970), the interest rate should be the monetary policy instrument when the LM curve is unstable and money supply should be the instrument when the IS curve is unstable. If the choice of the monetary policy instrument is inappropriate, monetary policy will increase the costs of stabilisation. Since instability in the demand for money is a major factor contributing to instability in the LM curve, it is important to test for the stability of the demand for money. Many industrial countries switched to the interest rate as the monetary policy instrument when their money demand functions became unstable following the financial reforms from the second half of the 1980s. Many developing countries have also abandoned controlling money supply and have been using the rate of interest as the monetary policy instrument, even though there is no significant evidence that their demand for money functions have become unstable.

Second, estimates of the demand for money are useful for understanding the limits to non-inflationary seigniorage revenue and for the formulation of monetary policy targets. Third, the unit roots and cointegration literature has had a significant impact on modelling dynamic economic relationships and especially on modelling the demand for money. Thus, there have been a large number of empirical studies, in both industrial and developing economies, to re-estimate the demand for money and to investigate, afresh, its stability; see Sriram (1999) for a survey. Fourth, Perron's (1989) influential work that the standard unit roots tests lose power if the variables undergo structural changes led to developments in testing for unit roots and estimation of the cointegrating equations.

In comparison to the large amount of empirical research on the demand for money in other countries, there is only a handful of such studies for Fiji. Furthermore, the existing studies are difficult to obtain and seem to have limitations, both in the specification and estimation of the relationship. Therefore, in this paper we review two recent empirical studies on Fiji with a view to providing a starting point for further work and highlight key issues for further investigation. However, the scope of our paper is not exhaustive because a single paper is not adequate to examine and resolve all the relevant issues.

Empirical studies for Fiji

Earlier papers on the demand for money in Fiji, such as International Monetary Fund (1982), Luckett (1984), and an unpublished study by Joynson (1997), are hard to obtain. A brief review of these studies is in Katafono (2001). To conserve space, these papers are not reviewed here. Recently, Jayaraman and Ward (2000) estimated a quarterly model of the demand for broad money and found that it was stable over 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. The income elasticity was insignificant with a t-ratio of 1.33 and the sign of the real interest rate elasticity, measuring the return on quasi money, was 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-money, dominates its negative effect, the cost of holding narrow money. Katafono (2001) pointed out that Jayaraman and Ward's finding, using quarterly GDP data, that the demand for money is stable is of little use for policy because it is not possible to forecast the quarterly 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 estimate of income elasticity is insignificant.

Given such limitations in the earlier work, the recent study by Katafono (2001) is a significant contribution and is, in our view, a good starting point for further work in Fiji and other Pacific island countries. Therefore, in the rest of this section we review her work in some detail. Katafono estimated the demand functions for narrow, quasi and broad money in Fiji for the period 1975 to 1999 utilising annual data. However, the demand for narrow money (or M1 as it is commonly known) received relatively more attention in her study. Since our objective is to estimate and analyse the demand for M1, we only review this part of her work.

Katafono used a standard specification of the demand for money (M1, henceforth M), in a semi-logarithmic form

$$\ln\left(\frac{M_t}{P_t}\right) = a_0 + a_1 \ln Y_t + a_2 SVR_t + a_3 TBR_t + a_4 REER_t + e_t \qquad (1)$$

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 e_t is an *iid* error term.[2]

After the unit root tests showed that these are all I(1) variables, Katafono conducted cointegration tests on these variables using the Johansen maximum likelihood method (JML) and found that there is one cointegrating vector. She 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 empirical work on the demand for money—she 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]

$$\ln\left(\frac{M_t}{P_t}\right) = -2.964 + 0.610\ln Y_t - 0.190\,SVR_t + 0.104\,TBR_t - 0.048\,REER_t + e_t \quad (2)$$

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 using a systems method, she estimated, with OLS, a variant of this equation, based on the well-known general to specific (GTS) approach of Hendry.[4]

In Katafono's two estimates there are some minor differences in the estimated long-run coefficients from GTS and JML. It would have been valuable if, for comparison, she had estimated a parsimonious dynamic adjustment equation using the lagged residuals from the cointegrating equation. The implied long-run relationship of her GTS estimates is[5]

$$\ln\left(\frac{M_t}{P_t}\right) = 0.511\ln Y_t + -0.104\,SVR_t + 0.004\,TBR_t + -0\,REER_t + e_t \quad (3)$$

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 argues that the sign of the coefficients of *SVR* should be positive. This is contrary to what she has found and also contrary to 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 be negative. However, the coefficients of *TBR* are positive in both equations, which may be due to collinearity between these two rates of interest. Perhaps Katafono should have re-estimated both equations after deleting *TBR* from the specifications. The major problem with her two estimates is that they imply an implausibly low income elasticity of about 0.5, contrary to her claim that it is close to unity. (See Jayaraman and Ward 2000:Table 1.4 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, Katafono's final estimate of the demand for money is found to be temporally unstable.

Unit root tests

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

$$\ln\left(\frac{M_t}{P_t}\right) = a_0 + a_1\ln Y_t + a_2 R_t + e_t \quad (4)$$

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_t is an iid error term. Our sample period extends from 1971 to 2002. Definitions of the variables and sources of data are given in the 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 from 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 hypothesis for the levels of the variables at the conventional 5 per cent or 10 per cent 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. ADF (WS) also shows that the unit root null hypothesis cannot be rejected for the levels of the variables. Similarly, the computed Elliott-Rothenberg-Stock test statistics (ERS)

Table 1 **Tests for unit roots: levels and first differences of variables with intercepts and linear trends**

Variable	*m*	ADF	ADF(WS)	PP	ERS
$\ln\left(M_t/P_t\right)$	[2,2,2,0]	–2.26	–0.73	–6.38	15.84
		(0.44)	(0.99)	(0.71)	(5.72)
$\Delta\ln\left(M_t/P_t\right)$	[1,1,1,0]	–5.17	–5.42	–39.20	1.87
		(0.00)	(0.00)	(0.00)	(2.97)
$\ln\left(Y_t/P_t\right)$	[2,3,3,1]	–1.44	–1.87	–16.80	21.28
		(0.81)	(0.73)	(0.13)	(5.72)
$\Delta\ln\left(Y_t/P_t\right)$	[2,2,2,0]	–4.30	–2.40	–37.76	2.63
		(0.00)	(0.06)	(0.00)	(0.97)
R_t	[2,5,5,2]	–1.79	–0.90	–2.26	69.34
		(0.61)	(0.98)	(0.96)	(5.72)
ΔR_t	[1,1,1,0]	–4.49	–4.75	–32.15	1.56
		(0.00)	(0.00)	(0.00)	(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 per cent critical values are shown in parenthesis. 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.

are more than the 5 per cent critical values, implying that all the levels of the variables are non-stationary. However, the p values for the first differences of these variables are significant at the 5 per cent level and reject the unit root null hypothesis. The computed test statistics for ERS are also below the 5 per cent critical values. Therefore, these variables are I(1) in levels and I(0) in their first differences.

Empirical estimates: general to specific approach

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

In GTS, first, a 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 general specification is reduced to a parsimonious dynamic adjustment equation using the variable deletion tests, by ensuring that the summary statistics do not become significant and reject the null hypothesis that the residuals satisfy the underlying classical assumptions.[7]

There is 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

$$\ln\left(\frac{M_t}{P_t}\right) = a_0 + a_1 \ln Y_t + a_2 R_t + e_t \qquad (5)$$

can be written as[8]

$$\Delta\ln\left(\frac{M_t}{P_t}\right) = b_0 + b_1 \ln\left(\frac{M_{t-1}}{P_{t-1}}\right) + b_2 \ln Y_{t-1} + b_3 R_{t-1} + b_4 \Delta\ln\left(\frac{M_{t-1}}{P_{t-1}}\right) + m_t \qquad (6)$$

Although Equation 6 seems simple, it is computationally demanding because the general dynamic specification will include many lagged values of the relevant variables. Furthermore, there are no clearcut guidelines on how to reduce the long lag structure to arrive at a manageable parsimonious equation for estimation. The general dynamic version of Equation 6 can be specified as

$$\Delta\ln\left(\frac{M_t}{P_t}\right) = b_0 + b_1 \ln\left(\frac{M_{t-1}}{P_{t-1}}\right) + b_2 \ln Y_{t-1} + b_3 R_{t-1} + l_i \sum_{i=0}^{n} \Delta \ln Y_{t-i} + g_i \sum_{i=0}^{m} \Delta R_{t-i} + t_i \sum_{i=1}^{j} \Delta \ln\left(\frac{M_{t-1}}{P_{t-1}}\right) + m_t \qquad (7)$$

This specification retains the error correction part, given by the lagged levels of the variables, and the equilibrium long-run coefficients are given by β_0, β_1, β_2 and β_3. If the three I(1) level variables are cointegrated, since their first differences are stationary, the error term μ_t will be I(0) and satisfy the standard classical assumptions. Therefore, OLS can be used to estimate Equation 7.

However, we 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 effect of political uncertainty on the demand for money. It is reasonable to expect that its coefficient would be positive because political uncertainty is likely to increase holdings of precautionary balances. Second, there were devaluations in 1987 and

Table 2 **GTS short-run adjustment equations**

	7a	7b	7c	7d	7e	7f
Intercept	–4.455	–3.966	–1.532	–2.147	–2.139	–2.047
	(–0.73)	(–0.53)	(–1.04)	(–5.87)*	(–5.98)*	(–5.34)*
Trend	–0.009	–0.008		–0.002		
	(–0.51)	(–0.36)		(–0.047)		
$\ln\left(M_{t-1}/P_{t-1}\right)$	–1.151	–1.205	–1.189	–1.199	–1.169	–1.109
	(–5.07)*	(–5.10)*	(–5.13)*	(–5.31)*	(–5.51)*	(–5.21)*
$\ln\left(Y_{t-1}/P_{t-1}\right)$	1.475	1.451	1.107	1.199	1.169	1.109
	(1.68)	(1.38)	(4.22)*	(5.31)*	(5.51)*	(5.21)*
R_{t-1}	–0.034	–0.037	–0.036	–0.037	–0.034	–0.031
	(–2.48)*	(–2.75)*	(–2.79)*	(–2.74)*	(–2.93)*	(–2.53)*
$\Delta\ln\left(Y_t/P_t\right)$	1.922	1.785	1.599	1.646	1.646	1.742
	(3.03)*	(2.65)*	(4.27)*	(3.78)*	(3.86)*	(4.71)*
$\Delta\ln\left(Y_{t-4}/P_{t-4}\right)$	0.816	0.838	0.809	0.813	0.832	0.802
	(2.03)**	(1.63)*	(1.68)	(2.12)*	(2.23)*	(1.69)**
ΔR_t	–0.052	–0.055	–0.051	–0.053	–0.049	–0.045
	(–2.33)*	(–2.70)*	(–2.91)*	(–2.54)*	(–2.64)*	(–2.68)*
COUP	0.314	0.322	0.280	0.296	0.265	0.247
	(2.61)*	(2.45)*	(3.79)*	(3.61)*	(5.41)*	(4.07)*
DEVDUM		–0.030	–0.031	–0.031	–0.031	
		(–2.27)	(2.36)*	(–1.08)	(–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
$c^2(sc)$	0.713	0.004	0.111	0.060	0.033	0.265
	(0.40)	(0.95)	(0.74)	(0.81)	(0.86)	(0.61)
$c^2(ff)$	4.578	4.329	3.553	3.553	3.331	3.461
	(0.03)	(0.04)	(0.06)	(0.06)	(0.07)	(0.06)
$c^2(n)$	0.040	0.088	0.087	0.107	0.079	0.178
	(0.98)	(0.96)	(0.96)	(0.95)	(0.96)	(0.92)
$c^2(hs)$	3.376	2.827	3.955	3.658	3.741	4.408
	(0.07)	(0.09)	(0.05)	(0.06)	(0.05)	(0.04)

Notes: t-ratios are in parenthesis. For the c^2 test statistics, p-values are in the parenthesis. * and ** signify 5 per cent and 10 per cent significance levels, respectively. In equations where $c^2(hs)$ is significant, we have used the Newey-West adjusted standard errors.

1998. Devaluations increase the prices of imported goods, although there should be a lag between the devaluation and the increase in the prices of imports. This is the well-known exchange rate pass-through effect. The effects of devaluations, therefore, would be immediate but transitory. Immediately after devaluations, there should be an increase in the purchase of imported goods, causing a shift from holding money to holding 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 (NBF) in 1996 might have caused a loss of confidence and a shift away from bank money in particular. These confidence loss effects seem to have persisted. 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.

A few parsimonious versions of Equation 7 are reported in Table 2. In Equation 7a, in the second column, all the summary statistics are satisfactory, except that the functional form misspecification $c^2(ff)$ (RESET) test is significant at the 5 per cent level, although not at 1 per cent level. The t-ratios (in parentheses below the coefficients) indicate that the estimated income elasticity is insignificant even at the 10 per cent 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 on the 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 the slightly higher side at 1.28. However, this is not unusual for developing economies (see Jayaraman and Ward 2000 for estimates of the income elasticities of some developing economies).

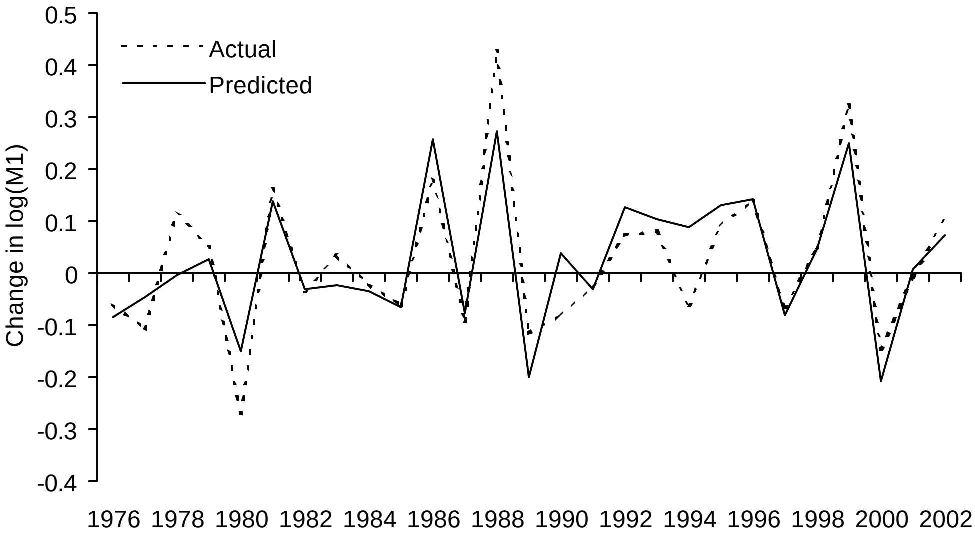
When this equation was tested for temporal stability with TIMVAR tests, the CUSUM test indicated instability from 1998 onwards, but the CUSUM SQUARES test showed that it is stable. To try to improve the summary statistics we added the dummy for the devaluations in 1987 and 1998. This did not improve the results and its coefficient was insignificant. We also added the dummy variable for the collapse of the National Bank of Fiji in 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 the estimate of the income elasticity declined marginally to 1.20 but is significant only at the 18.5 per cent 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.

The trend variable remained highly insignificant in both equations. Although it is essential to include a trend variable in VAR models, plots of real money, real output, and the rate of interest show that these variables are not strongly trended in Fiji. Therefore, we tested for the constraint that the coefficient of the trend variable is zero. The computed $c^2(1)$ test statistics is 0.16 and significant only at 69 per cent. Therefore, Equation 7c in Table 2 is estimated without the trend. The summary statistic improved except $c^2(2)_{hs}$ for heteroscedasticity, which is now significant at the 5 per cent level. Three changes are noteworthy. First, the estimate of income elasticity is almost unity. Second, the functional form misspecification $c^2(2)_{ff}$ statistic is insignificant at the 5 per cent 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 $c^2(1)$ test statistics for this constraint are 0.08 and 0.32 and significant at 78 per cent and 57 per cent respectively. Therefore, Equations 7d and 7e, that is, with and without trend, were estimated with the constraint of unit income elasticity. Both equations are well determined. None of the c^2 summary statistics are significant at the 5 per cent level and all other coefficients, except those of the dummy variable, are significant at the 5 per cent level. When these equations are subjected to the TIMVAR tests, the CUSUM and CUSUM SQUARES 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 have similar summary statistics and standard errors of estimates. The standard errors of estimates of our equations are similar to those reported by Katafono (2001). The constrained Equation 7e has the lowest standard error of estimates (SEE) of 0.08. 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 using Equation 7d are plotted in Figure 1. It can be seen that the fit is fairly good except for 1978, 1983, and from 1989 to 1995.[10]

Figure 1 **Actual and predicted values of Dln M from Equation 7d**



Source: Authors' calculations.

Cointegration and ECM estimates: Johansen method

The JML procedure in Microfit was used to test 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) variables selected were the intercept, the 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 selected VAR(1). We postponed the Granger causality tests until we tested for the number of the cointegrating vectors. The JML estimates implied that the null hypothesis of no cointegration could be rejected for VAR(1) but not for VAR(2). In VAR(1) the null hypothesis that the number of cointegrating vectors is zero (r = 0) was rejected by the trace test statistic only at the 10 per cent level. The computed value is 28.98 and the 10 per cent critical value is 28.78. The null that r = 1 is accepted by the eigenvalue and trace test statistics at the 95 per cent level. The single cointegrating vector, normalised on money, and obtained with JML is

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

We conducted weak and strong exogenity tests for the null hypothesis that money does not Granger-cause income and the rate of interest. The computed $c^2(2)$ test statistic for the weak exogenity test, with the p-value in parenthesis, is 6.04 (0.049). The corresponding strong exogenity test statistics $c^2(4)$ is 11.59 (0.021). In both cases, the null hypothesis can be accepted only at the 1 per cent 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 normalised on real money. The two key coefficients on income and the rate of interest have the expected signs and magnitudes. The estimated income elasticity of demand for money is almost unity at 1.13, in comparison to Katafono's estimate of about 0.7. The implied interest elasticity, at the mean interest rate of 6.97, is –0.286, which is also plausible. These elasticities are comparable to recent estimates by Pradhan and Subramanian (2003) for India 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 the 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 the following equations, t-values are given below the coefficients in parentheses and p-values are given in parentheses for the c^2 summary statistics. Significance at the 5 and 10 per cent levels is indicated with * and ** respectively.

$$\Delta \ln\left(\frac{M_t}{P_t}\right) = -2.952 - 0.002\,T - 1.079\,ECM_{t-1}$$

(–6.38) (–0.47) (–6.22)*

$$+1.136\,\Delta \ln Y_t - 1.246\,\Delta \ln Y_{t-1} - 0.827\,\Delta \ln Y_{t-2}$$

(–3.07)* (–2.86)* (–2.18)*

$$+0.826\,\Delta \ln Y_{t-4} - 0.039\,\Delta R_t + 0.035\,\Delta R_{t-1}$$

(2.84)* (2.62)* (3.08)*

$$+0.269\,COUP - 0.119\,DEVDUM$$

(4.36)* (-1.76)** (9)

$\bar{R}^2 = 0.825$, SEE = 0.061, Period: 1976-2002

$c^2(sc1) = 0.411\ (0.52)$, $c^2(ff) = 5.68^*\ (0.02)$

$c^2(n) = 0.684\ (0.71)$, $c^2(hs) = 0.780\ (0.38)$

The summary statistics of Equation 9 are good and a noteworthy feature of the equation is that it has a 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 $c^2(ff)$ test is significant at the 5 per cent level but not at the 1 per cent level. This is not unusual for dynamic equations because it is hard to claim that the complex nature of dynamic adjustments can be adequately captured with linear specifications and limited data. All coefficients are significant except that of the time trend. The combined devaluation and NBF dummy is significant at the 10 per cent level. When we tested separately for the significance of the devaluations in 1987 and 1998 and the failure of the NBF, the second devaluation seems to have had the largest impact. Therefore, the above equation was re-estimated with a dummy variable only for the second devaluation. Its coefficient is significant at the 10 per cent level. The functional form misspecification test statistic deteriorated somewhat but is still insignificant at the 1 per cent level. These estimates are given in Equation 10.

$$\Delta \ln\left(\frac{M_t}{P_t}\right) = -2.952 - 0.002T - 1.079\,ECM_{t-1}$$
$$(-6.38)\ (-0.475)\ (-6.22)^*$$

$$+1.136\,\Delta \ln Y_t - 1.125\,\Delta \ln Y_{t-1}$$
$$(-3.07)^* \qquad (-2.86)^*$$

$$-0.827\Delta \ln Y_{t-2} + 0.827\Delta \ln Y_{t-4} - 0.038\Delta R_t$$
$$(-2.18)^* \qquad (2.84)^* \qquad (2.62)^*$$

$$+0.035\Delta R_{t-1} + 0.269COUP - 0.119DEV2$$
$$(3.08)^* \qquad (4.36)^* \qquad (-1.75)^{**} \quad (10)$$

$\bar{R}^2 = 0.825$, SEE = 0.061, Period: 1976–2002

$c^2(sc1) = 0.411\ (0.52)$, $c^2(ff) = 5.68^*\ (0.02)$

$c^2(n) = 0.684\ (0.71)$, $c^2(hs) = 0.780\ (0.38)$

It is possible to reduce the number of estimated coefficients in Equation 9 and Equation 10 to increase the degrees of freedom. The positive coefficient of ΔY_t is close in value to the absolute values of the coefficient of ECM_{t-1} and $\Delta \ln Y_{t-1}$. Furthermore, the coefficients of ΔR_t and ΔR_{t-1} are close and opposite in sign. The coefficients of $\Delta \ln Y_{t-2}$ and $\Delta \ln Y_{t-4}$ are also close and opposite in sign. When these four restrictions are tested, the computed $c^2(4)$ test statistic (with the p-value in parenthesis) is 0.37 (0.985) and insignificant. Therefore, this constraint could not be rejected. The following ultra parsimonious Equation 11 is based on these restrictions

$$\Delta \ln\left(\frac{M_t}{P_t}\right) = -3.047 - 0.002T - 1.114\,ECM_{t-1}$$
$$(-10.94)^*(-0.706)\ (-11.05)^*$$

$$+1.114\Delta^2 Y_t - 0.820\left(\Delta \ln Y_{t-2} - \Delta \ln Y_{t-4}\right)$$
$$(11.05)^* \quad (-3.59)^*$$

$$-0.039\Delta^2 R_t + 0.269COUP - 0.119DEV2$$
$$(-6.43)^* \qquad (5.93)^* \qquad (-1.93)^{**} \quad (11)$$

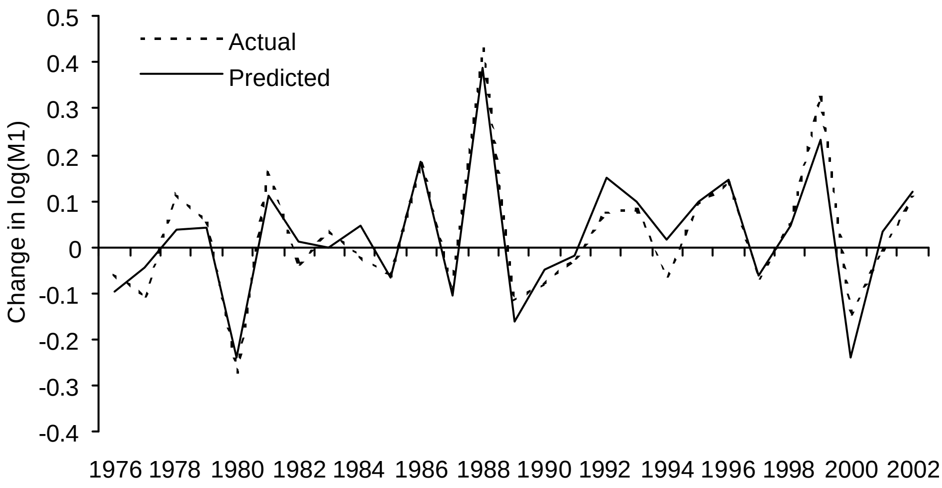
$\bar{R}^2 = 0.857$, SEE = 0.055, Period: 1976–2002

$c^2(sc1) = 0.169\ (0.68)$, $c^2(ff) = 3.94^*\ (0.05)$

$c^2(n) = 1.053\ (0.59)$, $c^2(hs) = 0.642\ (0.42)$

The summary statistics of this equation are impressive and the estimated coefficients are

Figure 2 **Actual and predicted values of Dln M from Equation 11**



Source: Authors' calculations.

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 available from the authors on request. The predicted and actual values from Equation 11 are plotted in Figure 2.

Conclusions

This article has surveyed earlier research on the demand for money in Fiji. It is noted that Katafono's (2001) work has much merit and is a good starting point for further work. However, while Katafono's study is relatively free of the weaknesses in earlier work, it is also found to be in need of improvement. Therefore, we have used two different time series methods of estimation of the demand for money in Fiji. The GTS and JML methods yielded similar cointegrating coefficients, although their dynamic adjustment lags are somewhat different. The estimated income and interest rate elasticities are found to 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 about the appropriateness of the 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 noted. First, several test statistics used are appropriate only for large samples.

Therefore, in further work it is important to make adjustments to minimise our finite sample biases. Such adjustments need considerable computational effort and thus they fell outside the scope of this paper. Second, we have ignored structural breaks and their implications for the unit root tests and estimation of the cointegrating relationships. 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, there are two ways of proceeding with estimation of 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 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 identify cointegrating relationships and therefore it is not useful for developing countries with limited 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 work. Consequently, before our findings are used for policy formulation in Fiji, we emphasise the need for further work based on more refined techniques and better insights into the theory of the demand for money. We hope that our work and methodology will be useful, together with Katafono's, as starting points for further research in this area in Fiji and in other Pacific island countries.

Appendix

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

P = GDP deflator. The 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 IMF, *International Financial Statistics* (2003 CD-ROM).

R = Nominal interest rate. 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 IMF, *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 to 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 Inclusion of the real rate of interest in the demand for narrow or broad money is difficult to justify because nominal rates of return on various liquid assets and their close substitutes are all equally affected by inflation, leaving the relative rates unchanged. Therefore, in several reputable works on the demand for money, and in textbook discussions, only nominal rates of interest are included (see Friedman 1969; Laidler 1969; Hendry and Ericsson 1991a, 1991b; Mishkin 2002). 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 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. International Monetary Fund (1982) and Jayaraman and Ward (2001) include the real rate of interest. A similar approach was used by Ahmed (2000) in estimating the demand for money in Bangladesh. Use of such a specification might be due to the mistaken notion that since the demand for money depends on real income it should also depend on the real rate of interest.

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 multi-collinearity between these two rates of return.

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

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 rate of return variable, for example, a weighted average of deposit rates in trading partner countries. We have ignored this variable because we consider that foreign exchange holdings are not a realistic option in Fiji.

7 A good exposition of GTS can be found in Charemza and Deadman (1997). The Davidson, Hendry, Srba and Yeo (1978) work on the consumption function for the United Kingdom is a classic paper on GTS. Subsequently, Hendry and Ericsson (1991a, 1991b) used this approach to re-estimate and test the money demand function for the United States of Friedman and Schwartz (1982).

8 This formulation is based on an exposition in Cuthbertson (1995).

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 adjusted $\bar{R}^2$ = 0.755 and SEE = 0.072.

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

12 While the theoretical developments in analysing structural changes are valuable, Maddala and Kim (1998) 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, an oil price shock) that occurred at time t. 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 1998:398).

These observations imply that perhaps testing for unit roots with *a priori* known dates, for example, 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.

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