# Estimating the Demand for Money in an Unstable Open Economy: The Case of the Fiji Islands

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ABSTRACT

*In this paper, we estimate Fiji's money demand function for the period 1971-2002 based on the bounds testing approach to cointegration, which is applicable irrespective of whether or not the underlying variables are non-stationary. We estimate models with and without a time trend and for lag lengths ranging from 1-3, but fail to find any evidence for a long-run relationship. Moreover, our structural break analysis suggests that the unstable nature of Fiji's money demand may be due to atypical events, such as coups; the implementation of policies, such as devaluations and value added tax; and the onset of trade liberalisation policies over the last two decades.*

## 1. INTRODUCTION

A simple money demand function implies that demand for money depends upon a variable that reflects the level of transactions in the economy such as real income or wealth, and another variable that reflects the opportunity cost of holding money such as the interest rate. Theoretically the relationship between income and money demand is expected to be positive, while the relationship between the interest rate and money demand is expected to be negative. These relationships are important for they are relevant in macroeconomic analysis, particularly in selecting appropriate monetary policy actions, where the stability of the money demand function is considered to be an important prerequisite for effective money targeting.

Moreover, a monetary policy aimed at restricting the supply of money facilitates demand management and is able to achieve price stability. The rate of growth in money supply should be in conformity with the desired rate of growth in output, thus allowing policy makers to maintain a control over price increases. It follows that a stable money demand function implies a stable money multiplier and, therefore, stability ensures that predicting the effect of a given money supply on aggregate money income is possible (see Pradhan and Subramanian, 2003). Hamori and Tokihisa (2001) note that the stability of the money demand function is important in order to give credence to the hypothesis that monetary policy matters, that is, money supply will have a certain amount of expected influence on real variables and that money supply control by the central bank is an effective macroeconomic policy.

The goal of this paper is to re-estimate the money demand function for the Fiji Islands. This is not the first attempt at undertaking this exercise for Fiji. Previous published studies that have estimated a money demand function for Fiji include Jayaraman and Ward (2000), Katafono (2001) and Rao and Singh (2005). Our main motivation for revisiting Fiji's money demand function is concern over weaknesses in the estimation techniques used in previous studies. A weakness common to the three studies mentioned above is that none of them consider structural break unit root tests. This is despite the fact that over the last three decades the Fijian economy has experienced several shocks, ranging from coups in 1987 and 2000 to devaluations in 1987 and 1998. Narayan (2004a,b, 2005a,b) shows that Fiji's macroeconomic variables are likely to be effected by structural changes. Such shocks are likely to impact the unit root properties of the commonly used variables, namely money demand, GDP and interest rates, in estimating the money demand function. That structural changes matter can be confirmed by simple plots of the natural log of the real money demand, natural log of real GDP and the nominal interest rate variables (see Figures 1-3).

**Figure 1: Plot of the natural logarithm of real money demand**

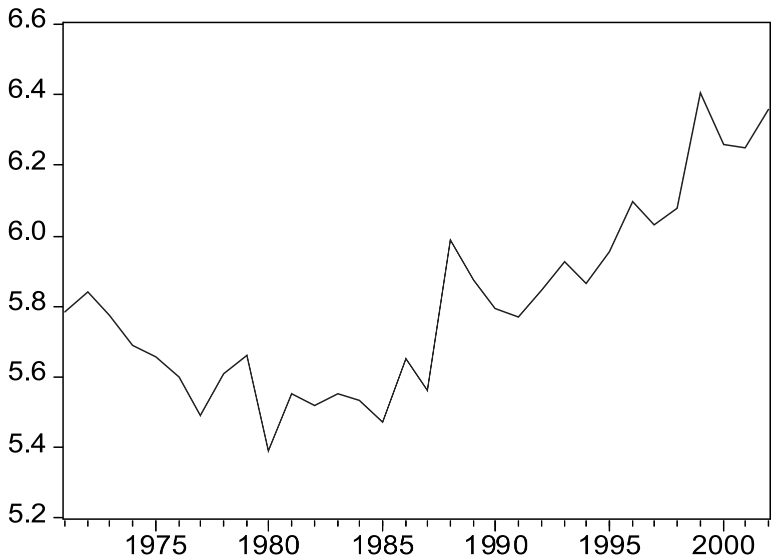


**Figure 2: Plot of the natural logarithm of real money demand**

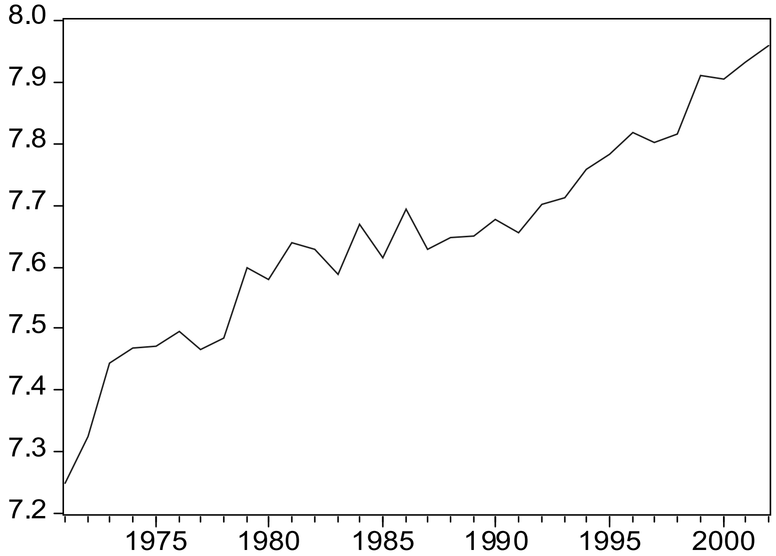
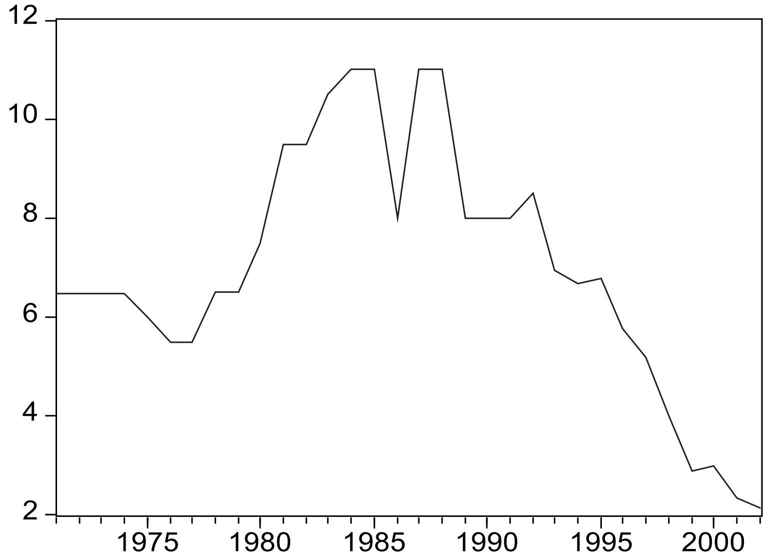



**Figure 3: Plot of nominal interest rate**




In testing the unit root properties of variables, which have implications for subsequent empirical analysis such as cointegration and stability tests, these studies fail to reject the unit root null hypothesis. Perron (1989) argues that if there is a structural break, the power to reject a unit root decreases when the stationary alternative is true and the structural break is ignored. Vogelsang and Perron (1998) show that the power of a unit root test with structural breaks is in the range 60-90 per cent under alternative scenarios.

Thus, it is fair to claim that that previous studies estimating Fiji's money demand function have failed to search fully for the integrational properties of the data series, which has implications for cointegration tests. Existing test for cointegration, such as the Engle and Granger (1987) and Johansen (1988) tests, require forcing variables to be integrated of order one. It follows that if any one of the proposed variables is stationary, the results from cointegration tests may be spurious.

Briefly foreshadowing our main results, applying the Lee and Strazicich (2003) endogenous one-break unit root test, we find that real money demand, real GDP and nominal interest rate are breakpoint stationary. Given that the variables appeared stationary, we applied the bounds testing approach to cointegration — a technique applicable irrespective of whether or not the variables are non-stationary — and could not find any evidence for a cointegrating relationship among the variables. That we fail to find any evidence of cointegration suggests that there is no stable long run money demand function for Fiji. In an attempt to verify the potential causes of this unstable money demand, we undertake a more sophisticated test, namely the Bai and Perron (1998) structural break test. This test provides evidence of three structural breaks in real money demand, real GDP and the nominal interest rate. We were able to associate these breaks to atypical events, such as coups; the implementation of policies, such as devaluations and the value added tax; and the onset of trade liberalisation policies in Fiji. Apart from fully modelling the demand for money for Fiji, another novel contribution of this paper is that it shows the importance of accounting for structural breaks in time series modelling, for structural breaks influence the integrational properties of the underlying variables and, as a result, allows practitioners to choose the most appropriate modelling techniques without compromising the quality of policy advice. Unfortunately, in the money demand literature appropriate attention has not been devoted to the issue of structural breaks, leading to erroneous policy implications, as in Rao and Singh (2005).

The balance of the paper is organised as follows. In section 2, we provide a brief overview of selected recent literature on developing countries. In section 3, we present the econometric methodology. In section 4, we present the model and discuss the results, while in the final section we conclude with policy implications.

## 2. An overview of recent literature on developing countries

Given the policy relevance of estimating the money demand function, the empirical literature is voluminous; thus, it is impossible to review all the studies here. Our aim in this section is only to review selected recent studies on money demand functions for developing countries, with particular emphasis on studies that have used the bounds testing approach to cointegration.

Siddiki (2000) uses the bounds testing approach to cointegration and estimates the money demand function for Bangladesh for the period 1975 to

1995. He augments the money demand function with the unofficial exchange rate. He finds that real money demand is cointegrated with income, the interest rate and the exchange rate, and that there is a stable long-run relationship among the variables.

Bahmani-Oskooee and Shin (2002) estimate Korea's money demand function using the bounds testing approach to cointegration. They find that while the variables included in the money demand function are cointegrated, the parameters are not stable. They find a structural break in 1997, associated with the Asian financial crisis and claim that this external shock may have led to the instability.

Bahmani-Oskooee and Rehman (2005) estimate money demand functions for 7 Asian countries using the bounds testing approach to cointegration. They use the CUSUM and CUSUMSQ tests to examine the stability of the money demand functions. They find that for India, Indonesia and Singapore, M1 is cointegrated with its determinants and the parameters are stable, while the M2 is cointegrated with its determinants for Pakistan, Philippines, Malaysia and Thailand with stable parameters.

Tang (2002) estimates Malaysia's money demand function using the bounds testing approach to cointegration. Instead of using the aggregate income variable, he disaggregates income into its components, namely expenditure on final consumption goods, expenditure on investment goods and exports. He finds a cointegrating relationship between money demand and its determinants; and using the CUSUM and CUSUMSQ tests he finds a stable money demand function for Malaysia.

Pradhan and Subramanian (2003) estimate the money demand function for India using the Gregory and Hansen (1996) residual based test for cointegration, which allows for a regime shift. They do not, however, find conclusive evidence of a stable long-run relationship between money demand and its determinants for India. Sriram (2002) uses the Johansen (1988) test for cointegration to estimate Malaysia's money demand. He augments the conventional money demand model with the inflation rate and finds evidence for cointegration and a stable money demand function.

Khalid (1999) examines the money demand function for the Philippines, Singapore and South Korea using the Johansen (1988) approach to cointegration. He finds evidence for cointegration among real money demand, real income and the nominal interest rate, but he does not conduct any analysis for the stability of the money demand function.

Weliwita and Ekanayake (1998) augment the conventional money demand function with the exchange rate and foreign interest rate variables for Sri Lanka. Using the Johansen (1988) test they find a cointegrating relationship between money demand and its determinants, but they do not test explicitly the money demand function for stability.

## 3. Econometric Methodology

### 3.1. Lagrange Multiplier structural break unit root test

Lee and Strazicich (2003) propose a one-break Lagrange multiplier (LM) unit root test, which has the advantage that it is unaffected by breaks under the null. The LM unit root test can be explained using the following data generating process:

$$y_t = \delta' Z_t + X_t \qquad\qquad X_t = \beta X_{t-1} + \varepsilon_t \tag{1}$$

Here, $Z_t$ consists of exogenous variables and is an error term that follows classical properties. The LM unit root test allows for structural breaks in the spirit of Perron (1989). In the case of the one-break unit root test, $Z_t = [1, t, D_{1t}, DT_{1t}]$, where $D_{1t} = 1$ for $t \geq T_B + 1$, and 0 otherwise, and $T_{B1}$ represents the break date. Lee and Strazicich (2003) use the following regression model to obtain the LM unit root test statistic:

$$\Delta y_t = \delta' \Delta Z_t + \phi \bar{S}_{t-1} + \mu_t \tag{2}$$

where $S_t = y_t - \hat{\psi}_x - Z_t \hat{\delta}_t$, $t = 2,...,T$; $\hat{\delta}$ are coefficients in the regression of $\Delta y_t$ on $\Delta Z_t$; $\hat{\psi}_x$ is given by $y_t - Z_t$; and $y_1$ and $Z_t$ represent the first observations of $y_y$ and $Z_t$ respectively. The LM test statistic is given by: $\tau$ = $t$-statistic for testing the unit root null hypothesis that $\phi$=0. The location of the structural break is determined by selecting all possible break points for the minimum $t$-statistic as follows:

$$Inf\,\tilde{\tau}\left(\bar{\lambda}_i\right) = \ln_{\lambda} f\,\tilde{\tau}\left(\lambda\right) \text{ , where } \lambda = T_B/T$$

The search is carried out over the trimming region (0.15$T$, 0.85$T$), where $T$ is sample size. Critical values for the test are tabulated in Lee and Strazicich (2003).

### 3.2 Cointegration

In this section, we discuss the bounds testing approach to cointegration between money demand, real GDP and nominal interest rate for Fiji. The bounds test is constructed using an autoregressive distributed lag model (see Pesaran et al, 2001); thus we begin with an unrestricted VAR in levels:

$$X_t = \alpha + \sum_{j=1}^{p} \Psi_j X_{t-j} + \varepsilon_t \tag{3}$$

where $X_t$ = [*MD, Y, R*]$^t$: *MD* denotes real money demand, *Y* denotes real GDP, and *R* denotes short-term nominal interest rates; α is a vector of constant terms, $\alpha = \left[\alpha_{MD}, \alpha_Y, \alpha_R\right]$ and $\psi_j$ is a matrix of VAR parameters for lag *j*; and the vector of error terms $\varepsilon_t = \left[\varepsilon_{MD,t}, \varepsilon_{Y,t}, \varepsilon_{R,t}\right] \sim IN(0, \Omega)$, where Ω is a positive definite.

We can now write Equation (4) in a vector error correction model (VECM) form as:

$$\Delta X_t = \alpha + \delta X_{t-1} + \sum_{j=1}^{p-1} \theta_j \Delta X_{t-j} + \varepsilon_t \tag{4}$$

where Δ = 1-*L*,

$$\theta_j = -\sum_{k=j+1}^{p} \psi_k$$

and δ is the long-run multiplier matrix given by

$$\delta = -\left(\mathrm{I} - \sum_{j=1}^{p} \psi_j\right)$$

where *I* is a 3 × 3 identity matrix and the diagonal elements of this matrix are left unrestricted, such that each of the variables in our model can be either I(0) or I(1). The bounds testing approach to cointegration is imposed by simply rewriting Equation (4) as follows:

$$\begin{aligned}\Delta \ln MD_t &= a_{0MD} + \sum_{i=1}^{n} b_{iMD} \Delta \ln MD_{t-i} + \sum_{i=0}^{n} c_{iMD} \Delta \ln Y_{t-i} \\ &+ \sum_{i=0}^{n} d_{iMD} \Delta R_{t-i} + \lambda_{1MD} \ln MD_{t-1} + \lambda_{2MD} \ln Y_{t-1} + \lambda_{3MD} R_{t-1} + \varepsilon_{1t}\end{aligned} \tag{5}$$

The bounds testing procedure examines the absence of any level relationship between $MD_t$ and its determinants through exclusion of the lagged levels variables $MD_{t-1}$, $Y_{t-1}$ and $R_{t-1}$ in equation (2). It follows then that our test for the absence of a conditional level relationship between has the following null hypothesis:

$$H_0 : \lambda_{1MD} = 0, \quad \lambda_{2MD} = 0, \quad \lambda_{3MD} = 0 \tag{6}$$

This hypothesis can be examined using the standard Wald or $F$ statistics. The $F$-test has a non-standard distribution which depends upon (i) whether variables included in the ARDL model are I(0) or I(1), (ii) the number of regressors, (iii) whether the ARDL model contains an intercept and/or a trend; and the sample size. Critical values are extracted from Narayan (2005c).

If the computed $F$ statistics fall outside the critical bounds, a conclusive decision can be made regarding cointegration without knowing the order of integration of the regressors. For instance, if the empirical analysis shows that the estimated $F$ statistic is higher than the upper bound of the critical values then the null hypothesis of no cointegration is rejected. Once a long-run relationship has been established, in the second stage a further two-step procedure to estimate the model is carried out. First, the orders of the lags in the ARDL model are selected using an appropriate lag selection criteria such as the Schwartz Bayesian Criteria (SBC) and, in the second step, the selected model is estimated by ordinary least squares.

### *3.3. Number and location of break points in a money demand function*

Conditional on finding all variables to be stationary, we proceed to estimating breakpoints for real money demand, real GDP and nominal interest rates using the Bai and Perron (1998) method. Bai and Perron recommend a multiple linear regression model with $m$ breaks ($m$+1 regimes). The model takes the following form:

$$y_t = x_t'\beta + z_t'\psi_j + \kappa_t, \qquad t = T_{j-1}+1,\ldots,T_j, \quad \text{for} \quad j = 1,\ldots,m+1 \tag{7}$$

Here $y_t$ is the dependent variable at time $t$, $x_t$ $(p \times 1)$ and $z_t$ $(q \times 1)$ are vectors of covariates, $\beta$ and $\psi_j$ are the corresponding vectors of coefficients, and $\kappa_t$ captures disturbance at time $t$. In Equation (7) the break points $(T_1,\ldots T_m)$ are explicitly treated as unknown. Our objective is to estimate the unknown regression coefficients together with the break points when $T$ observations of $(y_t, x_t, z_t)$ are available. Bai and Perron develop test statistics to identify multiple breaks. For each $m$-partition $(T_1,\ldots T_m)$, $\beta$ and $\psi_j$ are estimated by minimising the sum of squared residuals:

$$\sum_{i=1}^{m+1} \sum_{t=T_{i-1}+1}^{T_i} \left[ y_t - x_t'\beta - z_t'\psi_j \right] \tag{8}$$

Letting $\hat{\beta}\left(\{T_j\}\right)$ and $\hat{\beta}\left(\{T_j\}\right)$ denote the estimates based on the given $m$-partition (T1,...Tm) represented here as $\{T_j\}$, one needs only to substitute these into Equation (8) to get the estimated breakpoints. If we depict the resulting sum of squared residuals as $S_T,\ldots,(T_1,\ldots T_m)$ it follows the estimated break points $(T_1,...,T_m)$ will be such that $(T_1,...,T_m) = \arg\min_{T_1,...,T_m} S_T(T_1,...,T_m)$ (see Bai and Perron, 1998).

Bai and Perron recommend a suite of tests for (1) ascertaining the existence of one or more structural breaks in a series, and (2), conditional on finding that there are such breaks, determining the number and location of breaks. To test for the existence of one or more structural breaks, they advocate the $SupF_t(L)$ *F*-statistic and double maximum tests. The $SupF_t(L)$ *F*-statistic tests the null hypothesis of no structural breaks ($m=0$) against the alternative hypothesis that there are $m=k$ breaks. The procedure searches all possible break points and minimises the difference between the restricted and unrestricted sum of squares over all the potential breaks.

The double maximum test considers the null hypothesis of no structural breaks ($m=0$) against the alternative hypothesis of at least 1 through $M$ structural breaks. The double maximum test takes two forms, which Bai and Perron term $UD_{max}$ and $WD_{max}$. The $UD_{max}$ statistic is the maximum value of the $SupF_t(L)$ *F*-statistic where $L$ represents an upper bound on the possible number of breaks, while the $WD_{max}$ statistic weights the individual statistics so as to equalise the *p*-values across values of *m*.

We use the *LWZ* procedure to determine the optimal number and location of structural breaks. The procedure tests the null of *L* breaks against the alternative of (*L* + 1) breaks. For the model with *L* breaks, the estimated break points denoted by $\hat{T}_1,...,\hat{T}_m$ are obtained by global minimisation of the sum of squared residuals. Rejection in favour of a model with (*L* + 1) breaks occurs if the overall minimum value of the sum of squared residuals is sufficiently smaller than the sum of squared residuals from the *L* break model. The appropriate critical values are available in Bai and Perron (1998).[2]

Before implementing the Bai and Perron (1998) procedure, an initial trimming region must be specified to ensure that there is a reasonable number of degrees of freedom to calculate an initial error sum of squares. The trimming specification determines the maximum possible number of breaks and minimum regime size. We imposed trimming $\in = 0.15$ and allowed the system to search for a maximum of three breaks. Given our small sample size, we believe that 3 breaks are sufficient. The Bai and Perron (1998) procedure also corrects for serial correlation in the errors and different variances of residuals across segments through incorporating Andrews' (1991) robust standard errors. The structural break model that we employ here for each of the three variables utilises this correction.

## 4. Model and empirical results

### *4.1. Model*

Our proposed long-run model is based on the specification that money demand $(MD_t)$, proxied by M1 (currency in circulation and demand deposits),[3] is a linear function of real GDP $(Y_t)$ and nominal interest rate $(R_t)$, proxied by

1-3 years weighted average interest rate on time deposits. The model takes the following form:

$$\ln MD_t = \alpha_0 + \alpha_1 \ln Y_t + \alpha_2 R_t + \varepsilon_t \tag{9}$$

In addition to the model description above,  denotes the natural logarithm and ε is the iid error term. The dataset spans the period 1971 to 2002 and is obtained from Rao and Singh (2005), who provide a detailed description of their source in an appendix to their article.

*4.2. Unit root tests*

The Lee and Strazicich unit root test that accounts for one endogenous structural break is reported in Table 1. For the real money demand variable, we are able to reject the unit root null hypothesis at the 1 per cent level, since the estimated *t*-test statistic of -12.21 is smaller than the critical value of -5.11 at the 1 per cent level. For the real GDP and nominal interest rate variables, we are able to reject the unit root null hypothesis at the 5 per cent level. The obtained *t*-test statistics are -4.49 for real GDP and -5.04 for the nominal interest rate.

**Table 1: Lee and Strazicich (2003) *LM* one-break test results**

| | $M1$ | $Y$ | $R$ |
|---|---|---|---|
| Break Date | 1986 | 1983 | 1985 |
| $\alpha$ | -3.8428*** | -0.9642** | -1.2051*** |
| | (-12.2131) | (4.4891) | (-5.0366) |
| $t$ | -0.5583*** | 0.0870*** | -0.8237** |
| | (-11.6972) | (3.4051) | (-1.7576) |
| $D_{lt}$ | -0.5851*** | 0.1262*** | -5.3964*** |
| | (-8.1400) | (3.3192) | (-6.1881) |
| $DT_{lt}$ | 0.7139*** | -0.1163*** | 1.9008*** |
| | (12.6017) | (-3.2042) | (2.4699) |
| $k$ | 8 | 8 | 7 |
| Critical values for the *LM* test | | | |
| $\lambda$ | 1 per cent | 5 per cent | 1 per cent |
| 0.1 | -5.11 | -4.50 | -4.21 |
| 0.2 | -5.07 | -4.47 | -4.20 |
| 0.3 | -5.15 | -4.45 | -4.18 |
| 0.4 | -5.05 | -4.50 | -4.18 |
| 0.5 | -5.11 | -4.51 | -4.17 |

Notes: $\lambda_j$ denotes the location of breaks. * (**) *** denote statistical significance at the 10 per cent 5 per cent and 1 per cent levels respectively. α is the coefficient on the unit root parameter, *t* is the coefficient on the time trend, $D_{lt}$ is the coefficient on the break in the intercept, $D_{lt}$ is the coefficients on the break in the slope, and *k* is the optimal lag length.

Other results from this endogenous break unit root test worth highlighting in brief are (1) the time trend is statistically significant at the 1 per cent level for real money demand and real GDP and at the 5 per cent level for the nominal interest rate; and (2) for all the three variables, the break in the intercept and the break in the trend are statistically significant at the 1 per cent level.

*4.3. Cointegration test*

Given that all variables are found to be stationary we proceed to conducting the bounds test for cointegration; the results of which are reported in Table 2. The results are for a model with and for a model without a time trend. An important issue with the bounds test is the selection of the optimal lag length (Narayan and Smyth, 2006a). Given the implications for the finding of cointegration among the variables, it is important to get the correct lag length.

To avoid any uncertainty arising from lag length selection, we report the results for lag lengths ranging from 1-3. This seems to be more than sufficient, since Narayan (2005c) shows that for small sample sizes, a maximum of 2 lags should be used, a finding confirmed for a similar sample size to ours by Narayan and Narayan (2004, 2005).

**Table 2: Cointegration test results**

| | *90% level* | | *90% level* | | *90% level* | |
|---|---|---|---|---|---|---|
| Model | I(0) | I(1) | I(0) | I(1) | I(0) | I(1) |
| Without trend | 2.915 | 3.695 | 3.538 | 4.428 | 5.155 | 6.265 |
| With trend | 3.770 | 4.535 | 4.535 | 5.415 | 6.428 | 7.505 |

Calculated *F*-statistic

| *Lag length* | *No trend* | *Trend* |
|---|---|---|
| 1 | 1.8896 | 1.1988 |
| 2 | 1.5282 | 0.8747 |
| 3 | 2.5097 | 0.8484 |

Note: The critical values are obtained from Narayan (2005c).

Our results for the model without a time trend are 1.89, 1.53 and 2.51 for lag lengths 1-3 respectively. These F-test statistics, when compared with the exact small sample critical value from Narayan (2005c), are smaller than the upper bound critical value of 3.69 at the 10 per cent level, indicating that we do not find any evidence of a cointegrating relationship among the variables.

For the model with a time trend, the $F$-test statistics for lag lengths 1-3 are 1.20, 0.87 and 0.85, respectively. The upper bound critical value at the 10 per cent level is 4.55. Thus we find that the calculated $F$-test statistics are all lower than the 10 per cent level critical value. Hence we do not find any evidence for a cointegrating relationship between money demand, income and interest rate.

### 4.4. Multiple structural break test

Because all variables are stationary, the Bai and Perron (1998) test can be applied. The results from this test for real money demand, real GDP, and nominal interest rate are reported in Table 3. We find the $SupF_t(L)$ test to be significant at conventional levels for values of $L$ between one and three. The $UD_{max}$ and $WD_{max}$ test statistics are also both significant at the 1 per cent level. This implies that there is at least one structural break in each of the three variables used in the money demand function.

To select the optimal number of structural breaks we use the $LWZ$ procedure. Using this procedure, we find three statistically significant structural breaks for the real money demand, real GDP and nominal interest rate. In addition, we report the 95 per cent confidence interval for each break date for each of the variables. The importance of the confidence interval is crucial as it allows us to relate to incidence that occurred not only at the year of the break but the years surrounding the break.

Beginning with the real money demand series, we find structural breaks for the years 1987, 1994, and 1998. 1987 was the year of coups in Fiji. This break has a 95 per cent confidence interval spanning years 1984-1991. With the plummeting of Fiji's economy, the Reserve Bank devalued the Fiji dollar by 33 per cent in 1987: the aim was to boost Fiji's export competitiveness. Fears of insecurity due to a collapse of law and order triggered mass out migration of skilled professionals from Fiji, which instigated a trend of capital flight since then (Narayan and Smyth, 2003, 2005a, 2006b). In sum, the 1987 coups resulted in a major shock to money demand.

The second break year 1994 has a 95 per cent confidence interval spanning the years 1992-1998. This period was another of the most volatile ones with respect to Fiji's money demand. For instance, in 1992 the government introduced a value added tax policy, and in 1994/1995 the National Bank of Fiji scandal surfaced: over F$250 million, equivalent to 8 per cent of Fiji's GDP, was lost due to corruption. The third structural break of 1998 with a 95 per cent confidence interval spanning the years 1996 to 1999 is marked by both negative and positive shocks. While the year 1998 is one when the Reserve Bank devalued the Fiji dollar by 20 per cent, the early years of the confidence interval are one when the Fijian economy was still coming to terms with the collapse of its national bank. However, the year 1999 — the last year of the confidence interval — was one where democracy was returned to Fiji through the election of a government, which had boosted Fiji's democracy ratings and instilled some confidence in the economy.

**Table 3: Empirical results from the Bai and Perron (1998) test for multiple structural breaks in the money demand, GDP and interest rate for Fiji**

*Specifications*

$z_t\{1\}$ $q = 1$ $p = 0$ $h = 15$ $M = 5$

| | *M1* | *Y* | *R* |
|---|---|---|---|
| $SupF_t(1)$ | 21.8774***<br>(8.5800) | 0.1158<br>(12.2900) | 10.5818**<br>(8.5800 (5%)) |
| $SupF_t(2)$ | 97.0196***<br>(7.220) | 13.3304***<br>(9.3600) | 4.9453<br>(7.2200 (10%)) |
| $SupF_t(3)$ | 175.1858***<br>(5.9600) | 54.3987***<br>(7.6000) | 61.4268***<br>(7.6000 (1%)) |
| $SupF_t(2/1)$ | 60.3546***<br>(12.2900) | 5.2717<br>(7.0400) | 3.3832<br>(7.0400 (10%)) |
| $SupF_t(3/2)$ | 18.1686***<br>(14.8000) | 14.5267***<br>(13.8900) | 14.2314***<br>(11.1400 (1%)) |
| $UD_{max}$ | 283.2938***<br>(12.3700) | 54.3987***<br>(12.3700) | 61.4260***<br>(12.3700) |
| $WD_{max}$ | 175.1858***<br>(13.8300) | 87.9684***<br>(13.8300) | 99.3336***<br>(13.8300) |
| Number of breaks[2] | | | |
| Sequential | 3 | 2 | 1 |
| LWZ | 3 | 3 | 3 |
| BIC | 3 | 3 | 3 |
| Estimates with breaks | | | |
| $\hat{T}_1$ | 1987<br>(1984-1991) | 1978<br>(1976-1982) | 1980<br>(1978-1983) |
| $\hat{T}_2$ | 1994<br>(1992-1998) | 1993<br>(1991-1995) | 1988<br>(1987-1989) |
| $\hat{T}_3$ | 1998<br>(1996-1999) | 1998<br>(1995-1999) | 1997<br>(1996-1998) |
| $\hat{\psi}_1$ | 5.8669<br>(152.3857) | 7.6455<br>(542.6409) | 10.1875<br>(31.2991) |
| $\hat{\psi}_2$ | 6.0410<br>(118.6127) | 7.7962<br>(319.4555) | 7.0978<br>(23.1293) |
| $\hat{\psi}_3$ | 6.3166<br>(124.0222) | 7.9267<br>(290.5111) | 2.8740<br>(6.9806) |

Notes: [1] The $SupF_t(L)$ test and the reported standard errors and confidence intervals allow for the possibility of serial correlation in the disturbance. [2] We use a 5% size for the sequential test $SupF_t(L+1/L)$. In the parentheses are the *t*-statistics (robust to serial correlation) for $\hat{\psi}_j$ ($j$ = 1,…,2) and the 95% confidence intervals for $\hat{T}_i\,(i=1,2)$. **(*) indicate statistical significance at the 5% and 1% levels respectively.

Turning to the results for the real GDP and nominal interest rate series, the first break date and its associated confidence intervals for both series seem to be associated with the late 1980s and early 1990s recession is Fiji, sparked by the oil price crisis of the late 1980s. The second break date and confidence interval for real GDP falls in the period of major economic policy changes in Fiji, including trade liberalisation policies. In 1984, Fiji first signed an International Monetary Fund (IMF) structural adjustment package (Narayan and Smyth, 2005b). In the post-1984 period, Fiji moved towards an export-led strategy for growth. The coups had immediate and drastic adverse economic consequences for Fiji and the military-backed interim government saw greater industrialisation and an export-oriented model of development as central to Fiji's future economic development (Chandra, 1989: 170). The introduction of in-bond manufacturing sites and creation of tax-free zones generally stimulated manufacturing exports, in particular garment exports. The value of garment exports increased from F$8.8 million in 1987 to F$141 million in 1994 and to F$333 million by 2000 (Reserve Bank of Fiji, 2002: A42). In the early 1990s, garment export production made up over 55 per cent of manufacturing production, over 22 per cent of exports, and 37 per cent of employment in the manufacturing sector (Akram-Lodhi, 1992; Elek et *al.*, 1993). The growth in garment exports was facilitated by significant devaluation of the Fiji dollar following the 1987 coups and again in 1998.

The second break for the nominal interest rate falls in the period of political instability, which resulted in a 33 per cent devaluation of the Fiji dollar.[5] The third break for both the real GDP and nominal interest rate series marks the period of devaluation and the Asian financial crisis.

We take the analysis further by attempting to test the correlation and significance between money demand and the break dates relating to money demand together with some of the reasons for these structural breaks as discussed above. We run an ordinary least squares estimation of the relationship between money demand and the three structural breaks. These break dates are specified as dummy variables with the year of the break taking the value 1 and the rest of the years taking the value 0.

We consider additional dummy variables, such as those relating to the coups, the introduction of the VAT policy and the onset of trade liberalisation policies. For the coup variable, we set a dummy variable which takes the value 1 in years 1987, 1988 and 2000, while the rest of the years take the value 0. For the VAT dummy variable, years prior to 1992 take the value 0 while years from 1992 take the value 1.

Similarly, we construct a dummy variable for the onset of trade liberalisation by denoting the years prior to 1984 with values of 0 while years from 1984 were denoted with the value 1. In addition, as is conventional in estimating money demand functions, we include the real income and nominal interest rate variables. Notice that since the levels of these variables are stationary, we do not change the form of these variables.

The results are reported in Table 4. Panel A includes the OLS estimates together with the *t*-statistics, while Panel B includes the statistics on goodness of fit and diagnostics. We find that all coefficients are statistically significant at the 10 per cent level or better. More importantly, our dummy variables constructed on the basis of the structural breaks are all significant. The break dates exert a negative effect on money demand. The dummies relating to trade liberalisation, VAT and coups have had a positive impact on money demand. The message emerging from this OLS analysis is that these events in Fiji over the last couple of decades have led to the instability of money demand. Finally, we notice that the income elasticity is negatively related to money demand in the short-run. There are two possible explanations for this. First, Choi and Oh (2003) show that including financial innovation type variables increases income elasticity of demand. Our model does not include any financial innovation type variables due to data limitations, thus it is likely that including financial innovation type variables would increase the income elasticity. Second, in Fiji the climate for business activity, and investments in general, has been weak due to a sustained period of political instability beginning in 1987. In such an environment, where business activity is risky, the negative income elasticity in the short-run is not entirely surprising.

**Table 4: OLS estimates of short-run determinants of money demand for Fiji**

Short-run impacts

| *Regressors* | *Coefficients* | *t-statistics* |
|---|---|---|
| Constant | 6.7783*** | 2.7281 |
| $\ln Y$ | -0.1007** | -1.7431 |
| $R$ | -0.0559*** | -4.2441 |
| D1987 | -0.2667* | -1.6483 |
| D1994 | -0.1391*** | -3.4274 |
| D1998 | -0.0687** | -1.7431 |
| Tradelib | 0.2064** | 2.2925 |
| VAT | 0.1747** | 1.8905 |
| Coups | 0.2257* | 1.4577 |

Goodness of fit and diagnostics

| | *Statistics* | *Probability* |
|---|---|---|
| $R^2$ | 0.8429 | - |
| $\bar{R}^2$ | 0.7882 | - |
| $\chi^2_{NORM}(2)$ | 0.5022 | 0.7779 |
| $\chi^2_{SERIAL}(2)$ | 0.5981 | 0.5589 |
| $\chi^2_{ARCH}(1)$ | 0.0432 | 0.8367 |
| $\chi^2_{HETERO}$ | 0.7978 | 0.6323 |
| $\chi^2_{RESET}(1)$ | 0.0222 | 0.8829 |

Note: * (**) *** denote significance at the 10 per cent, 5 per cent and 1 per cent levels, respectively.

Previous studies have supported the idea that money demand may become unstable when subjected to domestic and/or international shocks. For instance, Chaisrisawatsuk et *al*. (2004: 19) argue: 'Once viewed as a pillar of macroeconomic models, the reputation of the demand function for real money balances has plummeted since the early 1980s. This development may be attributed to the destabilising effects of financial innovation and deregulatory measures in many countries'. Pradhan and Subramanian (2003) find evidence that financial deregulation and innovation did affect the stability of India's money demand. Sriram (2002) also shows that instability of money demand for Malaysia may be due to recessions/recovery and policy measures taken in response to these events to influence the financial system and exchange rates.

## 5. Conclusions

Estimating the money demand function is a crucial empirical exercise from a policy point of view. The idea is simple but carries substantial motivation. If the variables making up the money demand function are found to cointegrate and there is a stable long-run relationship then this is a justification for central banks to target money supply as a monetary policy tool. However, if the variables are not cointegrated it implies, de facto, that there is no stable long-run money demand function. In this case, it is justifiable to abandon targeting money supply, and thus there is a challenge for central banks to adopt an alternative monetary policy stance.

In the case of Fiji, we found an additional motivation for re-estimating the money demand function. Previous studies (Jayaraman and Ward, 2000; Katafano, 2001; and Rao and Singh, 2001) have all found a stable long-run money demand function. It was hard to reconcile the fact that a stable money demand function exists in an unstable economy such as Fiji, which has been intermittently shocked by coups and devaluations over the last couple of decades. The main weakness of previous studies on Fiji was that they failed to model correctly the integrational properties of the variables, which have serious implications for cointegration and stability tests given the prerequisite in most tests that all variables should be integrated of order one. Following the work of Perron (1989) and Perron and Vogelsang (1998), it is now well known that standard tests that do not account for structural breaks have very low power to reject the null hypothesis; thus, researchers almost always find a series to be non-stationary.

We achieved our goal in three steps. In the first step, we subjected the real money demand, real GDP and nominal interest rate variables to the Lee and Strazicich (2003) endogenous one-break unit root test. We found that we were able to reject the null hypothesis of a unit root for real money demand and real GDP at the 1 per cent level and for nominal interest rate at the 5 per cent. Given that the integrational properties are not an issue with the bounds testing approach to cointegration, and that small sample size critical values for this test are provided in Narayan (2005c), we tested for cointegration using this test. We estimated models with and without a time trend and for lag lengths ranging from 1-3, but failed to find any evidence for cointegration; thus, we reject the findings from previous studies that there is a long-run stable money demand function for Fiji. Katafono (2001) is perhaps an exception; however, she does not conduct any formal test for stability.

We took the analysis further by investigating multiple structural breaks through the use of a sophisticated technique suggested by Bai and Perron (1998). We found that each of the three series had undergone three structural changes. The idea here was to investigate possible reasons for the structural instability. We concentrated more on the structural changes revealed for the real money demand variable, and found that the break dates related to: atypical events, such as coups; the implementation of policies, such as devaluations and the value added tax; and the onset of trade liberalisation policies in Fiji.

We believe that the Reserve Bank of Fiji is justified in abandoning monetary aggregate targeting in favour of its current strategy of inflation rate targeting.[6] However, whether or not inflation targeting is a prudent option for a small open economy, such as Fiji, is beyond the scope of this paper and should form a fruitful research agenda.

In closing, we do not consider our study as the final chapter of research on Fiji's money demand function. Rather, we believe that this study provides a solid foundation for both policy makers and applied researchers to expand work on Fiji's money demand function. This exercise is worthy because of its policy relevance. In the light of this, one avenue for future work could be a nonlinear cointegration analysis and/or a fractional cointegration analysis in order to confirm the robustness of our results. Moreover, we show that structural breaks are a crucial consideration in conducting unit root tests. Structural breaks influence the integrational properties of the underlying variables and as a result it allows practitioners to choose the most appropriate modelling techniques without compromising the quality of policy advice. Unfortunately, in the money demand literature appropriate attention has not been devoted to the issue of structural breaks, leading to erroneous policy implications as in Rao and Singh (2005). We hope that the work conducted in this paper will be a source of motivation for future studies.

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## Endnotes

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2. The asymptotic theory for this test is strictly only valid for the case of non-trending data. With trending data there are different asymptotic distributions. However, Bai and Perron (2003, p. 14) state: 'the asymptotic distributions in the two cases are fairly similar, especially in the tail where critical values are obtained. Hence, one can safely use the same critical values. Using simulations we found the size distortions to be minor'.

3. The literature correctly suggests that the use of narrow money is more suitable in the case of developing countries because it better reflects the weak banking system and low financial sector development. Our initial aim was to estimate two models, one with narrow money and one with broad money. However, due to inconsistent data, we could not estimate a model with broad money.

4. We also estimated a model using the inflation rate but did not find any evidence of a long-run relationship. Our findings are, thus, contrary to those for other developing countries; see, for instance, Sriram (2002).

5. For a recent macroeconomic impact analysis of coups in Fiji, see Narayan and Prasad (2007).

6. Until 1989, the Reserve Bank of Fiji (RBF) heavily regulated the financial system, and the RBF exercised controls on the quantity of money supplied. The RBF used changes in statutory reserve deposits of commercial banks to influence the broad money and credit aggregates and ultimately inflation (RBF, 2004). However, in the post-1989 period, the RBF's monetary policy stance changed to interest rate targeting. The interest rate targeting operates through the control of the level of short-term interest rates through open market operations (OMO). Under the OMO, the RBF notes are sold in order to raise the interest rate and bought if the intention is to reduce the interest rate.

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