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Abstract

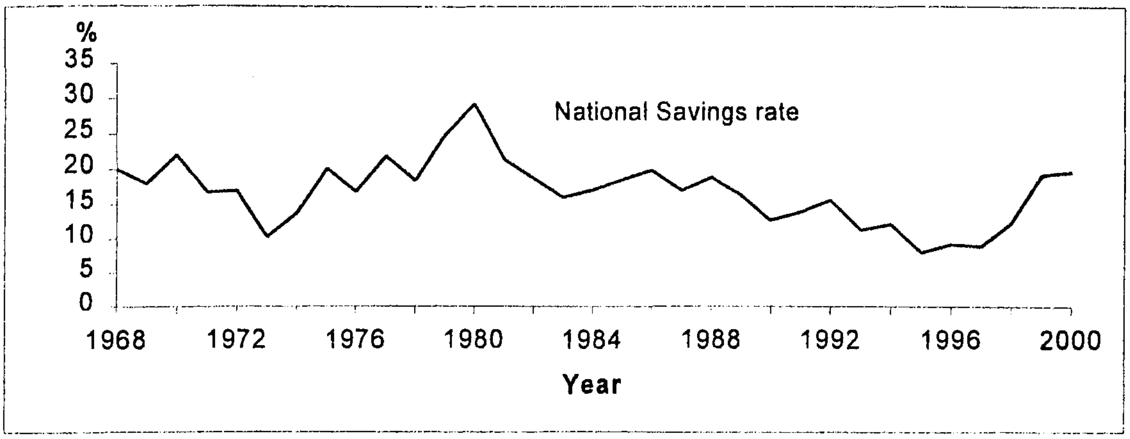
National savings play an important role in the economic development of many developing countries, especially if capital markets are weak. In this paper we investigate the determinants of national savings for Fiji. We use the recently developed autoregressive distributed lag modelling approach, shown to provide robust estimates in small samples, to model Fiji's savings behaviour. Our results indicate that the life cycle hypothesis helps explain Fiji's savings behaviour. The key finding is that economic growth has the biggest impact on savings rate, suggesting that savings will increase with an increase in economic growth.

Introduction

National savings play an important role in the economic development of many developing countries with limited avenues for financing investment projects. These countries, characterised by weak capital markets, are forced to rely heavily on domestic savings to finance such development projects. It is the shortage of national savings that is one of the most critical constraints on economic growth in most developing countries (Krieckhaaus, 2002). High national savings is likely to stimulate national investment, which in turn will provide the basis for more rapid economic growth (Nurkse, 1953).

Fiji Islands is a small isolated country with a weak capital market and is dependent mainly on national savings for financing investment in the country. It has a population of around 800,000 and is regarded as a lower-middle income country with per capita income of US$2110. Fiji's savings rate, however, has been declining throughout the 1980s and most of the 1990s and started to rise in the last four years (Figure 1). Total national savings as a percentage of gross national product has averaged only 17% in the 1968 to 2000 period. Given the importance of savings and the apparent low and downward trend over the last two decades, this study attempts to examine the factors that have affected savings in Fiji.

Figure 1: Savings Rate in Fiji, 1968 to 2000



Source: International Financial Statistics, 2001.

There are two other reasons for undertaking the present study. First, with the exception of a few recent studies (Agrawal, 2001; Sarantis and Stewart, 2001) on savings behaviour in developing countries, many studies (Khan *et al.* 1994; Cook, 1995; Cardenas and Escobar, 1998; Hussain and Brookins, 2001, amongst others) are open to criticism on econometric grounds for they ignore the fact that variables used in modelling saving rates are likely to be non-stationary and potentially integrated. It follows then that inferences made concerning long-run elasticities are potentially flawed and misleading, as noted by Yule (1926) and Granger and Newbold (1974).[1] Given this limitation in the literature on the determinants of savings rate, this study attempts to

[1] Yule (1926) suggested that regression based on non-stationary series are known as 'nonsense' regression while Granger and Newfold (1974) termed this problem 'spurious regression'.

correct for spurious regression by using cointegration analysis within the recently developed autoregressive distributed lag (ARDL) framework (Pesaran and Shin 1998 and Pesaran, *et al.* 2001), allowing estimates to be made of both short-run and long-run relationships.

Second, although studies on savings for groups of countries and regions have proliferated (see, *inter alia*, Giovannini, 1983; Edward, 1996; Callen and Thimann, 1997, Dayal-Gulhati and Thimaun, 1997; Hussain and Brookins, 2001), the focus has seldom been on a specific country. While this lacuna may be due to the lack of appropriate time series data on developing countries, no such problem exists for Fiji. The other issue is that given the use of cross-sectional data analysis in many studies, one cannot properly infer from observations across countries at a point in time what might happen in a country over time. A time series study overcomes this problem, providing a stronger basis for informed policy making.

The aim of this paper is to delineate the short- and long-run relationships between savings, real interest rate, income, current account deficits and age dependency ratio in Fiji using cointegration and error correction models over the period 1968-2000.

The balance of the article is as follows: the next section discusses the theoretical framework and provides a brief review of the literature; this is followed by a discussion of the econometric methodology to be applied, while the penultimate section presents the empirical results followed by the conclusion.

Theoretical Framework

Following on from various studies on determinants of savings (Modigliani, 1963; Giovannini, 1983; Doshi, 1994; Edwards, 1996; and Agrawal, 2001) the Modigliani-Brumberg-Ando 'Life Cycle' hypothesis is applied to examine the savings pattern in Fiji. The 'Life cycle' hypothesis contends that current savings decisions of households are a consequence of an act to distribute their lifetime consumption evenly over their lives so as to maintain the same lifestyle during retirement (Modigliani 1970). This hypothesis implies that savings rate depends on growth of income and/or the age structure of the population.

With respect to the growth rate of income, the life cycle theory postulates a positive relationship between savings rate and income growth rate, since higher income growth makes the young richer than the old; hence the young will be saving more than the old will be dis-saving. Recent empirical studies show this relationship to be true for some developing countries (Agrawal 2001). However, it is also possible that the effect of the rate of economic growth will be positive only to the extent that households, on average, accumulate wealth when they are young in order to dispose of this wealth when they are old. It follows then that growth rates may adversely affect savings rate in countries where households spend before they earn by borrowing. Furthermore, the rate of growth will have little or no effect on national savings if the age dependency ratio – defined as the population younger than 15 years of age plus population over 65 years old as a proportion of working age population - is high. Studies by Carroll (1992) and Carroll and Weil (1994) justifiably argue that a negative relationship between the two variables is also possible. Their argument is based on the notion that, other things being equal, an exogenous increase in growth level may make consumers feel wealthier and thus increase their consumption levels, leading to a lower savings ratio. Against this background, the expected sign of the savings growth rate is ambiguous.

The age structure of the population is also seen as an important determinant of savings in the life cycle model since a declining share of dependent population to working population is believed to correlate with higher savings rates. However, the existing empirical evidence on this relationship is mixed. For instance, Leff (1969) showed that this is true for both developed and the developing countries, while Ram (1982), Doshi (1994) and Kelly (1988), found no or very weak evidence of this relationship between savings rate and the age dependency ratio in developing countries. Agrawal (2001) found a significant negative relationship between savings and the age dependency ratio for three of the seven Asian countries.

Apart from the life cycle hypothesis, following Agarwal (2001), Hussain and Brookins (2001) and others, we employ current account deficits and real interest rates as additional explanatory variables. The relationship between national savings and current account deficits as a share of gross national product is clear cut - high foreign savings leads to a fall in national savings, since they are substitutes.

The real interest rate, defined as the difference between the nominal interest rate and the expected inflation rate, is the reward for postponing present consumption into the future. The net impact of the real interest rate on savings is unclear both theoretically and empirically, though the usual presumption is that the total effect is positive. One of the main reasons for the ambiguity follows from the distinction between the resulting income and substitution effects from an interest rate change (see Agrawal, 2001; Giovannini, 1983, 1985). The overall impact of interest rates on savings could be positive or negative depending on which effect is larger. Hence, the direction of the relationship is an empirical issue depending upon the relative strength of income and substitution effects. If the substitution effect outweighs the income effect savings would rise with an increase in real interest rate. Fry (1980, 1995) found that interest rates have a positive but small impact on savings while Giovannini (1983, 1985) found no statistically significant relationship between savings and interest rates. Agrawal (2001) found mixed results on this relationship for four Asian countries.[2]

Model specification

Based on the theory reviewed in the previous section, we posit the following specification to determine the national savings rate for Fiji.

$$S_t = \beta_0 + \beta_1 GR_t + \beta_2 IR_t + \beta_3 CAD_t + \beta_4 DEP_t + \varepsilon_t \qquad (1)$$

where S_t is the savings rate which is calculated as a ratio of gross national savings to gross national product; GR_t is the growth rate of per capita

[2] Taiwan's savings were significantly related to interest rate elasticities (-0.190), whereas savings from Malaysia, Korea and India were unaffected by real interest rates.

income; IR_t is the real interest rate[3]; CAD_t is the current account deficit as a proportion of GNP; and DEP_t is the share of dependents (below 15 and above 65) to the working population – the age dependency ratio.

The expected sign for β_3 and β_4 is negative, while the expected sign for β_1 and β_2 is ambiguous. We also considered the possibility of using the logarithmic version of the above variables to easily allow interpretation in terms of elasticities. However, this was impossible given that the current account deficit and growth rate of real per capita income take negative values.

Methodology

In the last 10-15 years, a number of tests for cointegration have been proposed in the literature. The two most commonly used cointegration techniques are the Engle Granger (1987) two-step residual-based procedure for testing the null of no-cointegration and Johansen's (1988,

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[3] The real interest rate is obtained using the formula: $R = [(1 + 1)/(1 + INF^{e})] - 1$, where R and I are the real and nominal interest rates on the yearly bank deposits and INF^e is the inflation rate calculated using CPI.

1991) system-based reduced rank regression approach. A common feature of these methods is their emphasis on cases in which the underlying variables are integrated of order one. Put differently, as a first step in estimation, the order of integration is of essence. This inevitably involves a certain degree of pre-testing, and thus introduces uncertainty into the analysis of levels relationships (Pesaran *et al.* 1996; Pesaran *et al.* 2001; see also Cavanagh *et al.* 1995).

The methodology used here is based on the recently developed autoregressive distributed lag (ARDL) framework (Pesaran and Shin, 1995, 1999; Pesaran *et al.* 1996; Pesaran, 1997; Pesaran *et al.* 1998) which does not involve pre-testing variables, thereby obviating uncertainty.[4] Put differently, the ARDL approach to testing for the existence of a relationship between variables in levels is applicable irrespective of whether the underlying regressors are purely $I(0)$, purely $I(1)$ or mutually cointegrated. The statistic underlying the procedure is the Wald or F-statistic in a generalised Dickey-Fuller type regression, which

[4] The pre-testing is particularly problematic in the unit-root-cointegration literature where the power of the unit root tests is typically very low, and there is a switch in the distribution function of the test statistics as one or more roots of the x_t process approach unity (Pesaran, 1997: 184).

is used to test the significance of lagged levels of the variables under consideration in a conditional unrestricted equilibrium correction model (ECM) (Pesaran *et al.* 2001, pp. 1)

Amongst other advantages, the ARDL method of cointegration analysis is unbiased and efficient. This is because it performs well in small samples, such as the present study. One can also estimate the long-run and short-run components of the model simultaneously, removing problems associated with omitted variables and autocorrelations. Finally, the ARDL method can distinguish dependent and explanatory variables. In what follows, the methodology is detailed.

The augmented ARDL $(p, q_1, q_2, \ldots, q_k)$ model can be written as follows (Pesaran and Pesaran, 1997, pp. 397-9).[5]

$$\Omega(L,p)y_t = \alpha_0 + \sum_{i=1}^{k} \beta_i(L,q_i)x_{it} + \delta' w_t + \mu_t \qquad (2)$$

where

$$\Omega(L,p) = 1 - \Omega_1\delta_1 L^1 - \Omega_2\delta_2 L^2 - \ldots - \Omega_p L^p, \qquad (3)$$

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[5] The model is extended and elaborated in detail in Pesaran and Shin, 1998 and Pesaran *et al.* 2001.

$$\beta_i(L,q_i)=\beta_{io}+\beta_{i1}L+\beta_{i2}L^2+\ldots+\beta_{iqi}L^{qi},\ i=1,2,\ldots,k, \qquad (4)$$

y_t is the dependent variable; α_o is a constant; L is a lag operator such that $Ly_t = y_{t-1}$; and w_t is a $s\times 1$ vector of deterministic variables such as the seasonal dummies, time trends, or exogenous variables with fixed lags. The x_{it} in Equation (2) is the i independent variable where i=1,2,…,k. In the long-run, we have $y_t = y_{t-1} = \ldots = y_{t-p}$; $x_{it} = x_{i,t-1} = \ldots = x_{i,t-q}$ where $x_{i,t-q}$ denotes the q^{th} lag of the i^{th} variable.

The long-run equation with respect to the constant term can be written as follows:

$$y=\alpha_o+\sum_{i=1}^{k}\beta_i x_i + \delta' w_t + v_t \qquad \Omega=\frac{\alpha_0}{\Omega(1,p)} \qquad (5)$$

The long-run coefficient for a response of y_t to a unit change in x_{it} are estimated by:

$$\beta_i=\frac{\hat{\beta}_i(1,\hat{q}_i)}{\Omega(1,\hat{p})}=\frac{\hat{\beta}_{i0}+\hat{\beta}_{i1}+\ldots+\hat{\beta}_{i\hat{q}}}{1-\hat{\Omega}_1-\hat{\Omega}_2-\ldots-\hat{\Omega}_{\hat{p}}}, \qquad i=1,2,\ldots,k \qquad (6)$$

where $\hat{p}$ and $\hat{q}_i$, $i = 1,2,...,k$ are the selected (estimated) values of p and q_i, $i = 1,2,...k$. Similarly, the long-run coefficients associated with the deterministic/exogenous variables with fixed lags are estimated by

$$\delta' = \frac{\hat{\delta}(\hat{p},\hat{q}_1,\hat{q}_2,...,\hat{q}_k)}{1-\hat{\Omega}_1-\hat{\Omega}_2-...-\hat{\Omega}_p}, \tag{7}$$

where $\hat{\delta}(\hat{p},\hat{q}_1,\hat{q}_2,...,\hat{q}_k)$ denotes the ordinary least squares estimate of δ in Equation (2) – the selected ARDL model.

The error correction (EC) representation of the $\mathrm{ARDL}(\hat{p},\hat{q}_1.\hat{q}_{2,...,}\hat{q}_k)$ model can be obtained by writing Equation (2) in terms of the lagged levels and the first differences of $y_t, x_{1t}, x_{2t,...,}x_{kt}$ and w_t:

$$\begin{aligned}\Delta y_t = \Delta\alpha_0 - \sum_{j=1}^{\hat{p}-1}\Omega_j^*\Delta y_{t-j} + \sum_{i=1}^{k}\beta_{i0}\Delta x_{it} - \sum_{i=1}^{k}\sum_{j=1}^{\hat{q}_t-1}\beta_{ij}^*\Delta x_{i,t-j} + \delta'\Delta w_t \\ -\Omega(1,\hat{p})ECM_{t-1} + \mu_t\end{aligned} \tag{8}$$

where ECM_t is the correction term defined by

$$ECM_t = y_t - \hat{\alpha} - \sum_{i=1}^{k}\hat{\beta}_i x_{it} - \delta' w_t \tag{9}$$

where Δ is the first difference operator; Ω_j^*, β_{ij}^* and δ' are the coefficients relating to the short-run dynamics of the model's convergence to equilibrium while $\Omega(1, \hat{p})$ measures the speed of adjustment.

A two step procedure is used in estimating the long-run relationship: an initial investigation of the existence of a long-run relationship among the variables in Equation 10 is preceded by an estimation of the short-run and long-run parameters (using Equation 1). This estimation is only possible if the long-run relationship is established in the first step.

Suppose that with respect to our model (Equation 1), we find that there is a long-run relationship among S_t, DEP_t, CAD_t, IR_t and GR_t. Without having any prior information about the direction of the long-run relationship among the variables, unrestricted error correction (EC) regressions are estimated, considering each of the variables in turn as a dependent variable; for example,

$$\begin{aligned}\Delta S_t &= a_{0S} + \sum_{i=1}^{n} b_{iS}\Delta S_{t-i} + \sum_{i=0}^{n} c_{iS}\Delta GR_{t-i} + + \sum_{i=0}^{n} d_{iS}\Delta IR_{t-i} \\ &+ \sum_{i=0}^{n} e_{iS}\Delta CAD_{t-i} + \sum_{i=0}^{n} f_{iS}\Delta DEP_{t-i} + \lambda_{1S}S_{t-1} + \lambda_{2S}GR_{t-1} + \lambda_{3S}IR_{t-1} \\ &+ \lambda_{4S}CAD_{t-1} + \lambda_{5S}DEP_{t-1} + \varepsilon_{1t}\end{aligned} \quad (10)$$

The F tests are used for testing the existence of long-run relationships. When long-run relationships exists, the F test indicates which variable should be normalised. The null hypothesis for no cointegration amongst the variables in Equation 10a is $(H_0 : \lambda_{1S} = \lambda_{2S} = \lambda_{3S} = \lambda_{4S} = \lambda_{5S} = 0)$ against the alternative hypothesis $(H_1 : \lambda_{1S} \neq \lambda_{2S} \neq \lambda_{3S} \neq \lambda_{4S} \neq \lambda_{5S} \neq 0)$. This can also be denoted as follows: $(F_S | GR, IR, CAD, DEP)$.

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 and (iii) whether the ARDL model contains an intercept and/or a trend. Two sets of critical values (CVs) are reported by Pesaran and Pesaran (1997) (repeated in Pesaran *et al.* 2001). Since these two sets of critical values provide critical value bounds for all classification of the regressors into purely $I(1)$, purely $I(0)$ or mutually cointegrated – the ARDL approach is also referred to as a bounds testing procedure (Pesaran *et al.* 2001).

If the computed F statistics falls outside the critical bounds, a conclusive decision can be made regarding cointegration without the need for knowing the order of integration of the regressors. For instance, if the empirical analysis shows that the estimated $F_S(.)$ is higher than the upper bound of the CV then the null hypothesis of no cointegration is rejected. In case that the computed F statistic falls inside the upper and lower bounds, a conclusive inference cannot be made without knowing the order of integration of the underlying regressors.[6]

Given that a long-run relationship exists, a further two step procedure to estimate the model is undertaken. The orders of the lags in the ARDL model are selected by either the Akaike Information criterion (AIC) or the Schwartz Bayesian criterion (SC).

Interpretation of the results

In the first step of estimating equation (1), equation (10) is estimated to examine the long-run relationship. Since the observations are annual, for the maximum order of lags in the ARDL we choose 2 and carry out the estimation over the period 1970-2000. The calculated F-statistics $F_S(.)$ is 4.053 which is higher than the upper bound critical value 4.049 at the 5 percent level, implying that the null hypothesis of no cointegration cannot be accepted and that there is indeed a cointegration relationship amongst the variables in equation (10).

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[6] For a 'unique and stable long-run' relationship $F_S($. should to be greater than the upper bound of the CV and $F_{GR}(.), F_{CAD}(.), F_{DEP}($. and $F_{DEP}(.)$ should be lower than the lower bound of the CV. In this relationship, S is the dependent variable GR, IR, CA and DEP are the explanatory variables.

Having found a long-run relationship, equation (1) is estimated using the following ARDL (m,n,p,q,r) model:

$$S_t = \alpha_0 + \sum_{i=1}^{m} \alpha_1 S_{t-i} + \sum_{i=0}^{n} \alpha_2 GR_{t-i} + \sum_{i=0}^{p} \alpha_3 IR_{t-i} + \sum_{i=0}^{q} \alpha_4 CAD_{t-i} + \sum_{i=0}^{r} \alpha_5 DEP_{t-i} + U_t \tag{11}$$

A maximum of 2 lags are used, i.e. $i = 2$ and the above model is selected using the Schwartz Bayesian criterion (SBC). The empirical results of the long run and the short run savings models for Fiji, obtained by normalising on the savings rate, are presented in Tables 2 and 3. All the coefficients have the expected signs in the long run.

In the long run, the growth rate of per capita income and the savings rate are positively related (Table 1) - consistent with some of the previous findings for developing countries (see Agrawal, 2001). The income growth elasticity of savings calculated at the mean value suggests that a 1% increase in growth rate leads to a 0.5% increase in savings. Growth in real income per capita in Fiji rose by an average of around 3% over the 1968-2000 period and has accounted for around 1.5% increase in the savings rate in the same period. The feeble income growth over the years has contributed to low savings growth in Fiji. Of all the factors, however, growth rate has the largest impact on national savings.

In the short run again growth rate exerts a positive impact on savings rate. However, the impact is smaller than in the long run (Table 2) – a 1% increase in growth rate leads to a 0.07% increase in savings rate which is statistically significant at the 5% level.

The current account deficit (CAD) appeared with the correct sign and is statistically significant at the 1% level in both the short and long run. The CAD elasticity of savings calculated at the mean value suggests that a 1% increase in CAD induced a 0.02% fall in savings rate; in the short run a 1% increase led to a 0.01% fall in savings rate.

Table 1: Econometric results for the long-run model, 1970-200

Variables (savings rate is the dependent variable)	Coefficient	t-statistics
Constant	0.2772*	1.4447
GR	0.0179*	1.4447
IR	-0.2204	0.5312
CAD	-1.5185***	3.6589
DR	-0.2220	0.8057

Note: * and *** indicate significance at the 20% and 1% levels respectively

Table 2: Econometric results for the short-run model, 1970-2000

Variables (savings rate is the dependent variable)	Coefficient	t-statistics
Constant	-0.00026	0.0457
ΔS_{t-1}	0.1073	0.6666
ΔGR_t	0.0024**	2.3902
ΔCAD_t	-0.8333***	4.7892
ΔIR_t	-0.0209*	1.3582
ΔIR_{t-1}	0.0246*	1.5191
ΔDR_{t-1}	0.1859	1.1902
EC_{t-1}	-0.1216**	2.4211

Notes: *(**)*** indicates significance at the 10%, 5% and 1% levels respectively.

The age dependency ratio produced mixed results. While in the long run it appeared with the theoretically correct sign – a 1% increase in the age dependency ratio leads to a 0.05% (elasticity calculated at mean value) fall in savings rate – in the short run a one period lagged effect was positive; a 1% increase in the age dependency ratio actually led to a 0.04% increase in savings rate. The age dependency ratio, however, had an insignificant impact – a result similar to the findings of Ram (1982), Kelly (1988), and Doshi (1994).

The effect of real interest rates on the savings rate has been a contentious one, with disagreements aplenty. Fry (1980, 1995) has found a positive relationship while others (Agarwal, 2001) have found both positive and negative relationships. Our results suggest that in the long run there is a negative but statistically insignificant relationship. In the short-run an increase in interest rates reduces the savings rate within that year, but a one-period lagged real interest rate significantly increases the savings rate. This may mean, *ceteris paribus*, that an increase in the interest rate induces higher spending in that year, indicating that the income effect dominates the substitution effect.

The error correction term EC_{t-1}, measures the speed of adjustment to restore equilibrium in the dynamic model, appeared with a negative sign and was statistically significant at the 5% level, ensuring that the series is non-explosive and that long-run equilibrium can be attained. The coefficient of –0.12, for instance, implies that a deviation from long-run savings rate in this period is corrected by about 12 percent in the next – an indication that once shocked convergence to equilibrium is slow.

Apart from the existence of a long-run relationship, our model is statistically well behaved. We applied a number of diagnostic tests to the error correction model (Table 3). There is no evidence of autocorrelation in the disturbance of the error term. The ARCH tests suggest the errors are homoskedastic and independent of the regressors. The model passes the Jarque-Bera normality tests suggesting that the errors are normally distributed. The RESET test indicates that the model is correctly specified, while the F-forecast tests indicates the predictive power/accuracy of the model. The adjusted R-squared of the model, 0.70, is reasonable – 70% of the variations in savings rate is explained by the regressors.

Table 3: Diagnostic tests for the ECM reported in Table 2

Diagnostics	
R^2	0.7535
$\bar{R}^2$	0.7026
σ	0.0308
$\chi^2_{Auto}(2)$	0.7701
$\chi^2_{Norm}(2)$	0.2078
$\chi^2_{WHITE}(14)$	10.8184
$\chi^2_{RESET}(2)$	0.3047
$F_{Forecast}(6,19)$	1.2893
$\chi^2_{ARCH}(2)$	5.0579

Where σ is the standard error of the regression; $\chi^2_{Auto}(2)$ is the Breusch-Godfrey LM test for autocorrelation; $\chi^2_{Norm}(2)$ is the Jarque-Bera normality test; $\chi^2_{RESET}(2)$ is the Ramsey test for omitted variables/finctional form; $\chi^2_{White}(20)$ is the White test for heteroscedasticity; $F_{Forecast}$ 6,19) is the Chow predictive failure test (when calculating this test, 1995 was chosen as the starting point for forecasting). Critical values for $\chi^2(2)$= 5.99.

Lastly, the stability of the regression coefficients is evaluated using the cumulative sum (CUSUM) and the cumulative sum of squares

(CUSUMSQ) of the recursive residual test for structural stability (Brown et al., 1975). The regression equation appears stable given that neither the CUSUM nor the CUSUMSQ test statistics exceed the bounds of the 5% level of significance.

Conclusions

In this paper we use the Pesaran *et al*, (2001) approach to cointegration to investigate the determinants of national savings in Fiji. The savings model is developed in line with the Life-Cycle hypothesis, which argues that the main determinants of savings rate are the growth rate of per capita real income and the age dependency ratio. The model was extended by including the current account deficit and the real interest rate variables. The results show that there is a long run equilibrium relationship between the above variables and they provide evidence that the life cycle model pertains to explaining savings behaviour in Fiji. Our key findings are as follows:

- In the short run and long run a 1% increase in growth rate increases savings by over 0.07% and 0.5% respectively;
- In the short run and long run a 1% increase in the current account deficit reduces savings rate by 0.01% and 0.02%, respectively;
- In the long run, the negative coefficient on the real interest rate implies that the income effect dominates the substitution effect, while

in the short-run the total effect of the real interest rate is positive, implying that the substitution effect dominates the income effect;

- We found the age dependency ratio negatively impacting the savings rate; however, this finding lacked statistical significance.

Of particular interest is the result on the growth rate of income per capita since it has the most influence on the savings rate. The result implies that, if a higher savings rate is to be achieved, policies have to be directed towards achieving higher economic growth. This seems an arduous task for policy makers in Fiji; Fiji's real GDP growth rate has been a mere 0.7% over the last decade. The abysmal growth performance is attributed to numerous factors, prominent amongst which is political instability. Fiji has experienced four coups since 1987 and Gounder (1999) provides empirical evidence demonstrating the negative relationship between coups and economic growth in Fiji. The last decade has also seen the emergence of land problems in Fiji. The majority of the agricultural land leases, which began expiring in 1997, have not been renewed; this is greatly affecting Fiji's sugar industry – the largest industry in terms of exports and employment. These problems, coupled with high rates of skilled labour emigration owing to sustained political instability and racial discrimination, have added to Fiji's woes; private investments averaged a mere 4% over the last decade. The fact that the legality of the current Fiji government is before the courts does not augur well for economic growth. In this light it is likely that national savings in Fiji will remain depressed.

Lastly, the contributions of this study should be noted. By using the recently developed autoregressive distributed lag (ARDL) approach to cointegration, we have corrected for the spurious regression problems inherent in many previous studies on the determinants of national savings. Given that Pesaran and Shin (1998) show that, under the ARDL framework, the OLS estimators of the short run parameters are $\sqrt{T}$-consistent, and the ARDL based estimators of the long run coefficients are super-consistent in small sample sizes, the results of our analysis are more robust than other studies. Furthermore, we have investigated an important issue facing a small island country, characterised by weak capital markets, where national savings is a crucial mechanism for providing development finance. In this light we believe that the results will be of interest to both policy makers in Fiji and academics.

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