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The Effects of Life Expectancy on Fiji's Output: A Time Series Approach from 1970 to 2002

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Abstract

Compared to several cross-country studies on the determinants of growth, time series approaches are relatively few and limited in scope. However, time series studies are useful for country-specific policies. But in the recent time series works, with a few exceptions, *ad hoc* specifications of output and growth equations are used. This paper examines the specification and estimation issues in the time series approach to the determinants of output. Our approach is used to measure the effects of health on the output of Fiji for the period 1970 to 2002.

JEL: N1, O1, O4;

KEYWORDS: The Solow Growth Model, Production Function, General to Specific Approach, Effects of Health on Output.

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I. INTRODUCTION

A stylized fact of growth and development accounting is that our knowledge of the determinants of growth is limited since factor inputs explain, at most, about half the observed variation in the growth rate. The remainder, the Solow residual, is attributed to the growth in technical progress (TP). However, the Solow residual is also a measure of our ignorance of the determinants of growth since it is not known what factors determine TP.[1]

Therefore, theoretical and empirical research has been directed at understanding the determinants of TP. The endogenous growth models of Romer (1986), Lucas (1988) and Barro (1991, 1999) have found that human capital, R&D and social infrastructure are also important determinants of growth. These factors are multi-dimensional, and empirical work based on the endogenous growth models have used different approaches to measure them.

An alternative line of research by Mankiw, Romer and Weil (1992) and Young (1995), extended the Solow model by broadening the measurement of inputs, e.g. the need to adjust employment for improvements in skills (human capital) and showed that factor accumulation can explain as much as 80% variation in the growth rate. A similar approach is advocated by Casseli (2004) to improve the measurement of capital. These are known as the chopping-off strategies because they chop-off the size of Solow residual. However, the chopping-off strategy is not beyond controversy; see for example Klenow Rodriguez-Clare (1997) and Hall and Jones (1999).

While the controversy on whether factor accumulation can explain

[1] The practice of treating the Solow residual as an estimate of TP and not as a measure of our ignorance can be best understood by noting that estimates of TP in Latin America were negative during 1980 to 1990. As Barro (1998) has observed, negative TPs are hard to understand because they imply forgetting technology. Therefore, it is appropriate to say that TP is a measure of our ignorance of positive and negative determinants of TP. Similarly it is hard to explain why TP has been low in the rapidly developing East Asian countries. However, it may be conjectured that such observed changes in TP might be actually due to shifts in the production function and/or inappropriate measurement of the factor inputs. Barro (1998) notes that, in an unpublished paper, Chang-Tai Hsieh showed that in Singapore the rate of growth of capital has been overestimated in official sources and if this is adjusted Singapore's TP was actually 0.022 and not −0.007 as found in Young (1995).

a major proportion of variation in growth is interesting and has theoretical implications on the relative merits of the neoclassical and endogenous growth models, more mixed pragmatic empirical approaches explore the determinants of output and growth from a policy perspective. As Barro (2000) has observed, from this empirical perspective, the new and old growth theories are more complementary than they are competing. In these approaches factors that contribute to output and growth are classified into two categories viz., those that affect output through their effects on the measurement of inputs and those that effect the technology intercept parameter by shifting the production function. Such shifts in the production function, however, increase the level of output permanently but not the growth rate, except in the transition period.[2] Since this transition period is long in actual calendar time, the distinction between level and growth effects seems to have been generally ignored in the empirical works. In this paper we shall refer to shifts in the production function as causing level effects and ignore the implied growth effects. To derive the transition dynamics of growth effects, it is necessary to derive the closed form solution for output and simulate the results. This is beyond the scope of our paper, but we shall briefly comment later when variables like human capital may also have permanent growth effects.

Variables like openness of the economy, institutional reforms, good governance and growth promoting policies etc., are likely to shift the production function. Improvement in education, health, work experience and imported capital etc., are likely to affect the measurement of inputs. This broad classification has implications for the specification and estimation of the output equation. Therefore, we first examine the specification issues and discuss whether health policy variables affect output through their effects by changing the measurement of labour, in the form of improved human capital, or directly by shifting the production function. Since education is considered to be an important determinant of human capital, it is difficult to develop a human capital index with health variables alone. Furthermore, it is also hard to separate the relative contributions of education and health to human capital formation. Nevertheless, human capital measures

[2] This is the same as whether, for example, an improvement in human capital is like Harrod neutral or Hicks neutral TP. Since there is a one to one correspondence between specifications in which a variable effects the measured inputs or the intercept of the production function, which of these alternative specifications is an empirical issue.

based on education alone are inadequate, because healthier workers are physically and mentally more energetic and productive. These issues are discussed in Section 2. Empirical results are presented in Section 3 and our conclusions are stated in Section 4.

II. SPECIFICATION

There are a few weaknesses in the recent country specific time series works. First they use *ad hoc* specifications. Often these works attempt to show that growth and one or two selected variables, say $X1$ and $X2$ are cointegrated and $X1$ and $X2$ Granger cause growth. Therefore a 10% increase say in $X1$ leads to a 2% improvement in the growth rate. In these works, there is also a confusion about the purpose of Granger causality tests, in spite of many warnings by Granger.[3] This point is especially emphasized by Stock and Watson (2003, p.449) with the observation that “While ‘Granger predictability’ is a more accurate term than ‘Granger causality’ the latter has become part of the jargon of econometrics.” Since the Granger tests are conceptually different from cause and effect tests, insights into the nature of cause and effect relationships should be based on the underlying economic theory. The endogenous growth theories are not developed after conducting a series of Granger causality tests.

A second weakness in such studies is that they are subject to serious omitted variables and misspecification biases. In both the cross section and time series growth and development accounting, specifications of growth equation are based on the neoclassical Solow growth model in which the rate of growth of output, *in the non-steady state,* depends on the rates of growth of inputs (capital and labour) and technology. Behind this growth equation is the neoclassical production function in which the level of output (Y) depends on capital stock (K), employment (L) and the stock of technology (A) and this can be stated as:

$$Y_t = F(K_t, L_t, A_t) \tag{1}$$

[3] If $Y = f(X)$, the Granger causality tests are essentially tests on whether X_t changes when there was disequilibrium in Y in period $t-1$. This is the weak exogeneity test. If ΔX_t is also not affected by ΔY_t, X is strongly Granger exogenous and ΔX_t may be included in the ARDL for ΔY_t. Granger causality tests are not, therefore, cause and effect tests.

In the time series work, the implication of this approach is that if these variables are I(1) in levels and are found to be cointegrated, the short run specification, based on error correction with a negative feedback (ECM), would have both the levels and the first differences of these key or conditioning variables. Therefore, estimating growth equations with only the rates of growth of the variables and ignoring their levels not only throws away valuable information but also causes other problems. Barro (1998) points out in the growth accounting context some additional problems with such specifications although he did not emphasis the need for retaining the levels of variables and the ECM formulation.

In the more recent cross-country development accounting works, which examine the effects of additional variables like education, health and institutional reforms etc., these conditioning variables are retained in the specification. Additional variables are augmented as shift variables into the neoclassical production function; see, for example, Bosworth and Collins (2003) and Bloom, Canning and Sevilla (2004). These requirements and implications for time series specification can be illustrated with a constant returns Cobb-Douglas production function with Hicks neutral technical progress as follows. However, our approach is different from the Bloom, Canning and Sevilla (BCS) approach and we shall shortly comment some differences.

$$Y_t = A_t K_t^{\alpha_1} L_t^{\alpha_2} \varepsilon_t \qquad \alpha_1 + \alpha_2 = 1 \tag{2}$$

where ε is an error term with the usual classical properties. The implied growth equation, although this is a misnomer, based on the cointegration and ECM specification, using for convenience the LSE-Hendry general to specific formulation (GETS) is:[4]

[4] Other alternatives can be used, but we think that it is convenient to use the GETS specification. Note that the standard criticism that in the LSE-Hendry approach I(0) first differences are mixed with (1) level variables and therefore the specification is not balanced, is not a valid criticism. Hendry repeatedly pointed out that if the I(1) variables are cointegrated then their linear combination is I(0). Furthermore, Banarjee *et.al* (1993) have shown that the GETS approach is as good as the estimates based on the fully modified OLS estimates. See also Hendry (1995) and Hendry and Doorink (1994).

$$\Delta lnY_t = -\lambda\Big[lnY_{t-1} - (\alpha_0 + \alpha_1 lnK_{t-1} + \alpha_2 lnL_{t-1})\Big] + \sum_{i=1}^{n1} \Delta lnK_{t-i} + \sum_{i=1}^{n2} \Delta lnL_{t-i} + \sum_{i=1}^{n3} \Delta lnY_{t-i} + \epsilon_t \quad (3)$$

where α_0 is lnA and ϵ is an error term with the usual classical properties. A similar specification was also recently used by Loening (2004) to estimate with time series data the effects of education on growth in Guatemala.

At this point it is important to note the conceptual differences between our and the BCS specification although these two are observationally somewhat equivalent. Our specification will be similar to the BCS specification if the lagged differences of the variables in (3) are replaced with their contemporaneous differences and their coefficients are constrained to be the same as the coefficients of the corresponding variable in the levels in the ECM part of equation (3). This difference is mainly due to the assumption that the production function (2) is an accounting identity in BCS, whereas it is a stochastic equation in our formulation. BCS derived a growth equation from their identity and assumed that the technology parameter A consists of two components– a systematic and a random component. The systematic component, since their production function is an identity, equals the difference between output and the inputs, weighted by their exponents. With the further assumption that output converges when there are random shocks, their specification of growth equation will be similar to the GETS specification except that the coefficients of the contemporaneous first differences and the corresponding level variables have the same coefficients.

From our perspective, the BCS constraints on the coefficients are restrictive, although in their empirical work with cross-country data these restrictions are validated. However, there are three problems in their specification. Firstly, as they have admitted, it is difficult to identify the coefficients of the technology parameters. Secondly, which is specific to time series regressions, the constraint that the coefficients of the levels and first differences of the variables may not be consistent with the underlying data generating process. Finally, it is hard to accept the assumption that the production function is an identity when its parameters are estimated. Nevertheless, the BCS approach is stimulating and worth testing if their coefficient constraints are valid in time series regressions. Since the BCS specification is nested in our specification, albeit in a mechanical sense, we shall estimate our preferred equation with the BCS constraints towards the end of this paper.

A final point to note is that while there is no distinction between the steady and non-steady states in the BCS cross-country regressions, the main objective of our equation (3) is to estimate the long run parameters of the production function. The implied growth rate captured by (3) is only for the short run. The long run effects of policy variables such as improvement in life expectancy, for example, are therefore only on the level of output and not on the growth rate. However, in the short run, an improvement in life expectancy may temporarily increase the growth rate. How long such temporary growth effects will continue can be computed by using the closed form solution for output, but, as stated earlier, this is outside the scope of the present paper.

In the rest of this paper, therefore, we shall continue with equation (3). Additional output determinants can be introduced in two ways. For illustration we shall use one additional variable viz., life expectancy (LE) since BCS have found that it is a good proxy for the effects of health on output and growth.[5] Our approach can be easily extended to include additional output determinants. Health can effect output either by increasing labour productivity, in the form of improved human capital, or directly by affecting the technology parameter A. This latter approach was used by BCS. Although it is hard to justify their assumption since it ignores the more direct effect of health through human capital formation, the BCS specification has some advantages. The specifications implied by both approaches can be developed as follows, but it would be easy to illustrate with the effects of education, rather than health, on human capital because it is difficult to develop an index of human capital with health related variables.

If education is introduced through its effects on improving human capital, the measurement of labour should be modified. This can be done by using the wage differences between workers with different levels of education as the weights to aggregate employment. For illustration, assume that there are two categories of workers, unskilled (no education) and skilled. Suppose skilled workers' wage rate is 50% higher.

[5] In our empirical work we have tried alternatives like the death rate, child mortality, real expenditure on health and the ratio of health expenditure to GDP. But, life expectancy is found to be the best of the lot.

If there are 50 workers in each category, unadjusted employment is 100. The adjusted measure of employment is $50 + 1.5 \times 50 = 125$. Thus the human capital component is 25. What this means is that the output produced by 50 unskilled and 50 skilled workers is 25% higher than the output produced by 100 unskilled workers. Therefore, the human capital index (HKI) is 1.25. It is hard to develop a similar index using health variables because data on the wage differences between healthy and unhealthy workers is not available besides the fact that it is also difficult to classify workers into different categories of health.

The growth equation, with ECM, ignoring the ARDL terms, will be:

$$\Delta lnY_t = -\lambda \Big\{ lnY_{t-1}[\alpha_0 + \alpha_1 lnK_{t-1} + \alpha_2 ln(L_{t-1} \times HKI_{t-1})] \Big\} \quad (4)$$

where HKI is an index of human capital component.

The approach of BCS is a variation of the old practice of introducing time trend to capture technical progress. If $A_t = A_0 EXP[rt]$, r is interpreted as the rate of technical progress and A_0 is the initial stock of knowledge. This approach indicates that there are some unknown highly trended factors that determine growth. The time trend can be replaced with myriad trended output determinants, except that unlike time they cannot increase indefinitely. Duraluf, Johnson and Temple (2004) have observed that the number of such potential determinants is as many as the number of countries in a cross-country study. The BCS specification is as follows:

$$Y_t = A_0 EXP[\Psi(HKI)] K_t^{\alpha_1} L_t^{\alpha_2} \qquad \alpha_1 + \alpha_2 = 1 \quad (5)$$

where Ψ could be linear or non-linear in variables such as life expectancy (LE) and education (ED). For simplicity we will assume a linear form. The growth equation, analogous to (4), without the ARDLs, will be:

$$\Delta lnY_t = -\lambda \Big\{ lnY_{t-1} - [a_0 + \pi_1 HKI_{t-1} + \alpha_1 lnK_{t-1} + \alpha_2 lnL_{t-1}] \Big\} \quad (6)$$

where HKI is an index of human capital. Although it is hard to justify, on theoretical grounds, that human capital forming variables like education (ED) and life expectancy (LE) would have their effects on output by affecting the technology parameter, the BCS specification is attractive for empirical work because any reasonable proxy can be used to capture the effects of HKI. There is also a one to one relationship between the underlying production functions of equations (4) and (6). For example, if Ψ is linear, as in (6), $\pi_1 = -[\alpha_2 ln(1/HKI) \div HKI] > 0$ since $ln(1/HKI) < 0$.[6] This makes clear that the BCS specification only shifts the production function and therefore variables like HKI have only permanent level and not permanent growth effect, unless it is reasonable to assume that HKI grows constantly at the rate of π_1 per period. To distinguish between whether a variable like HKI has a permanent growth effect or a level effect, it might be necessary to modify (6) so that HKI affects π_1. The following specification is an example:

$$\Delta lnY_t = -\lambda \Big\{ lnY_{t-1} - [a_0 + (\pi_1 \times HKI_{t-1})T + \alpha_1 lnK_{t-1} + \alpha_2 lnL_{t-1}] \Big\} \tag{7}$$

where T is time. It is difficult to discriminate between equations with level and growth effects because they seem to give similar results but computationally (7) was found to cause severe convergence problems. However, the estimates with level effects gave better results and therefore we shall ignore specifications based on (7).[7]

The BCS specification with level effects is especially useful when data are not available for a proper construction of HKI and as stated earlier this indeed is the case with health variables. Sometimes it is useful to mix both types of specifications when there is multi-

[6] π_1 can be solved by equating the two specifications of the production function, i.e., from:

$$A \times EXP[\pi_1 HKI]\ K^{\alpha_1} L^{\alpha_2} = A\ K^{\alpha_1} (L \times HKI)^{\beta_2}.$$

If HKI=1.15 and $\alpha_2 = 0.75$, then $\pi_1 \approx 0.1339$.

[7] When we experimented with the specification in (7), there have been severe convergence and multi colleniarity problems. In the end we have used the first principal component of TRADE, EDI and LEI and converted into an index number. Even then, we faced convergence problems. Therefore, we have experimented with plausible calibration techniques in which the adjustment coefficient λ was set to one and the share of profits to 0.33. This gave $\pi_1 = 0.00135$ and it was significant implying that the combined effect of the three variables over 33 years is to increase the growth rate permanantly to 0.00193. This implys that, on the average, growth rate permanantly increased by a negligible amount of 0.0000176 per year. Its $\bar{G}R^2$ is lower at 0.597 compared to 0.625 in our prefered equation. Much of the explanatory power seems to be due to the time trend. See equation (ii) in Table-1. These results are not reported to conserve space.

colleniarity between the variables. Therefore, we shall use both specifications, using life expectancy as a proxy for HKI.

III. EMPIRICAL RESULTS

All the variables in our empirical work are tested for unit roots with the ADF and KPSS tests and are found be, at least by one of these tests, as I(1) in levels and I(0) in first differences. Details can be obtained from the authors. Definitions of the variables, sources of data and the method of estimation of the capital stock with the perpetual inventory method are described in the data appendix.

We first estimated equation (3) with data, from 1970 to 2002, with the basic factor inputs and then with a time trend to see how inclusion of trend improves the fit. All equations are estimated with the two stage non-linear least squares instrumental variable method (2SNLLS-IV) to minimize endogeneity bias. Lagged values of the levels and first differences are used as instruments. The first two equations (i) and (ii) in Table-1 give, respectively, estimates of the basic growth equation with and without the trend. It can be seen that both equations are well determined, except that the adjustment coefficient λ in (i) is not significant. Their χ^2 summary statistics show that there is no serial correlation, functional form misspecification, non-normality of residuals and heteroscadisticity. The Saragan's χ^2 test indicates that the instruments are valid.[8] However, while equation (i), without trend, has a low explanatory power ($\bar{G}R^2 = 0.264$), equation (ii) with the trend shows a significant improvement with ($\bar{G}R^2 = 0.457$), almost double the explanatory power.[9]

In (i), the estimates of the share of profits and wages are significant and close to their stylized estimates of 0.3 and 0.7 respectively. The

[8] The Saragan test statistic is computed when there are overidentifying restrictions, with the null hypothesis that the selected instruments are exogenous i.e., they are uncorrelated with the error term. When the null is not rejected, it can be said that the chosen instruments are exogenous and valid. However, the Saragan test is appropriate for large samples whereas our sample size is modest.

[9] $\bar{G}R^2$ is a measure of goodness of fit of IV estimates and was developed by Pesaran and Smith (1994). It is a valid discriminator of models based IV method. Closeness between the $\bar{R}^2$ and $\bar{G}R^2$ is a rough indicator of how accurately the instruments predict the endogenous variables. In the extreme case where the instruments exactly predict the endogenous variables, both measures of goodness of fit will be identical.

constraint that there are constant returns could not be rejected by the Wald test. The computed test statistic, with p-value in square brackets, is $\chi^2 = .24902$ [0.618]. λ becomes significant when the constant returns constraint is imposed on (i) and the estimate is not reported to conserve space. There are no significant changes in the estimated coefficients. An interesting finding in (i) is that the 1987 coup variable decreased growth temporarily by about 0.07 points in 1987 and 1988, but this coefficient is insignificant.

Since the constant returns assumption is not rejected, equation (ii), with trend, is estimated with this constraint. Some notifiable changes are: the adjustment parameter λ has increased and is close to unity; share of profits declined from about 0.3 to 0.22 and the coup coefficient is significant at the 10% level, but its absolute size declined to 0.035. Thus equation (ii) is promising and indicates that augmenting the production function with additional trended explanatory variables improves the fit although we do not know what factors determine the effects of trend. When the actual and predicted growth rates are plotted 3 significant outliers in 1995 to 1997 are found. Therefore, equation (ii) is reestimated with an outlier dummy variable as equation (iii). It can be seen that adding the outlier dummy has significantly improved the fit with ($\bar{G}R^2 = 0.542$) and without any significant effect on the estimated coefficients. Since the Wald test could not reject the null that the coefficients of the dummies for coup and the outliers are opposite in sign and equal in absolute values, (iii) is actually estimated with this constraint. Both (ii) and (iii) imply that the unexplained trend rate of growth is modest at about half a percent per year. This small magnitude for the coefficient of trend partly indicates that perhaps there are no significant spillovers or endogenous growth effects in the production function for Fiji.

We have examined first the effects of life expectancy (LE) on output, using the specification implied by equation (4) in the text. An approximate proxy, life expectancy index (LEI), is developed using the 1970 life expectancy of 60.0488 years as the base, i.e. LE in 1970 is 1. Life expectancy in Fiji has gradually increased to about 69.5 years by 2002 or at the rate of about four months per year. LEI is used as a proxy human capital index to multiply the employment level. Estimates of the growth equation based on this adjustment are given as equation (iv) in Table-1. This equation is well determined and none of its χ^2 tests are significant. Its coefficient estimates are close to those in (iii) with the trend variable. The share of profits decreased marginally from 0.23 to 0.21 and similarly, its $\bar{G}R^2 = 0.531$ is only 0.01 less than that for equation (iii). We added the trend variable so see how inclusion of LE affects its coefficient. The coefficient of trend was significant but declined from about 0.005 in (ii) and (iii) to 0.0016. This equation is not reported to conserve space. On the basis of these results, It can be said that LEI is a potential determinant of output and growth in Fiji.

However, at this stage it is hard to draw any reliable conclusions about the quantitative effects of LEI because we have not yet included other potential determinants such as education, openness of trade, institutional reforms, good governance measures etc. It is difficult to include all such potential determinants and evaluate their relative contributions for two reasons. First, it is hard to obtain good time series indicators of these potential variables. Second, because of the likely multicollinearity between the variables it is also difficult to estimate their coefficients. Therefore, in what follows we shall make only a modest attempt to bring into our equations the role of openness of trade and education.[10] In the meantime, on the basis of the estimates in (iv), we can tentatively estimate a one year increase in LEI implies that it increases approximately by 0.017 points, which in turn implies that output will be 0.013 percent higher. In the short run, however, growth rate will increase temporarily by about 0.02 percent. If the 2005 GDP in Fiji is assumed to be \$3 billion, then real GDP will permanently increase by \$39.5 million.[11] This estimate, however, is likely to change when additional determinants of output are added and we shall now explore these changes.

Before additional determinants are added, it is worth reestimating equation (iv) using the BCS specification, but without their constraints on the coefficients of the levels and first differences of variables. The estimate with actual life expectancy (LE), using the BCS approach, are in equation (v) in Table-1. There is no point in estimating with the index LEI, since it is only a transformation of LE. This equation is well determined and none of its χ^2 summary statistics are significant. It is noteworthy that it has the highest $\bar{G}R^2 = 0.610$.

[10] We have experimented with the Freedom House measures of Political Rights and Civil Liberties, but their coefficients have been insignificant in our estimates.

[11] There are no official estimates of GDP for Fiji after 2002. In 2002 real GDP (in 1995 prices) was \$2,771 million. Assuming about 2.75% growth rate, GDP in 2005 will be \$3,009 million. The assumed 2.75% growth rate is rather optimistic, but serves to illustrate the computations.

TABLE-1
DETERMINANTS OF OUTPUT IN FIJI
2SNLLS-IV ESTIMATES

Variables	i	ii	iii	iv	v	vi
INTERCEPT	-4.0398 (-1.35)	-3.2457 (-24.25)	-3.2184 (30.98)	-3.2698 (-28.08)	-4.5288 (-17.82)	-4.2815 (-32.45)
λ	-0.5998 (-1.37)	-0.9844 (-5.96)	-1.1029 (-6.19)	-1.0101 (-5.71)	-1.2132 (-6.32)	-1.1953 (-6.372)
Trend		0.0050 (5.16)	0.0046 (6.49)			
LE_{t-1}					0.0184 (6.10)	0.0164 (8.24)
$\ln K_{t-1}$	0.3450 (2.41)	0.2230 (5.58)	0.2317 (7.10)	0.2086 (5.40)	0.1689 (4.73)	0.2086 (c)
$\ln L_{t-1}$	0.7676 (2.18)	0.7770 (c)	0.7683 (c)		0.8311 (c)	0.7914 (c)
$\ln(L_{t-1} \times LEI_{t-1})$				0.7914 (c)		
$\Delta \ln L_t$	-0.1932 (-0.23)	0.5024 (2.62)	0.4606 (2.63)		0.6337 (3.54)	0.5698 (3.42)
$\Delta \ln(L_t \times LEI_t)$				0.4925 (2.76)		
COUP	-0.0730 (-1.11)	-0.0353 (-1.83)**	-0.0383 (-2.67)	-0.0441 (-2.90)	-0.0415 (-2.84)	-0.0461 (-3.34)
OUTLIERS			0.0383 (c)	0.0441 (c)	0.0415 (c)	0.0461 (c)
$\overline{R}^2$	0.037	0.507	0.599	0.545	0.582	0.597
$\overline{GR}^2$	0.264	0.457	0.542	0.531	0.610	0.605
Saragan's χ^2	3.028 [0.220]	5.994 [0.112]	2.068 [0.558]	7.808 [0.253]	3.263 [0.860]	4.674 [0.792]
SEE	0.048	0.034	0.031	0.033	0.032	0.031
$\chi^2(sc)$	0.088 [0.767]	2.575 [0.109]	3.215 [0.073]	1.937 [0.164]	1.612 [0.204]	1.738 [0.187]
$\chi^2(ff)$	0.184 [0.668]	0.096 [0.756]	0.309 [0.578]	1.062 [0.303]	0.362 [0.547]	0.579 [0.447]
$\chi^2(n)$	0.511 [0.775]	1.540 [0.463]	0.839 [0.657]	0.187 [0.911]	1.074 [0.584]	0.555 [0.758]
$\chi^2(hs)$	0.046 [0.831]	0.161 [0.688]	0.034 [0.854]	2.391 [0.122]	2.285 [0.131]	2.954 [0.086]

Notes: t-ratios are in the parentheses below the coefficients; p-values are in the square brackets for the χ^2 tests; constrained estimates are denoted with (c).

However, it implies a somewhat lower share of about 17% for profits compared to the share of 0.21 in equation (iv). As it is, equation (v) with LE implies that a one year increase in life expectancy increases GDP permanently by about 0.18 percent. At the assumed value of \$3,000 million for 2005, this is about \$55.6 million, compared to about \$40 million implied by equation (iv). Because of the low profit share in (v), we have reestimated it with the constraint that profit share is the same as in equation (iv). The estimate is given as equation (vi) in Table-1. It can be seen that there is hardly any difference between (v) and (vi), except that the coefficient of LE has now declined to .0165, implying that a one year increase in life expectancy will raise output permanently by \$49 million, a figure in-between the estimates from equations (iv) and (v).

In spite of these encouraging preliminary results, it is important to postpone conclusions about the effects of health until the specification is extended to include other additional variables in the production function. For this purpose, it is found that neither of the two basic specifications based on equations (4) or (6) in the text are satisfactory. The BCS specification fails because of multi-collinearity between these variables. However, a mixture of both specifications turned out to be promising. In this hybrid approach, the effect of life expectancy is captured through its effect on the human capital component of employment i.e., as in equation (iv) in Table-1. The BCS specification is used to capture the effects of additional variables. When both trade and education variables are added to the production function, using the BCS specification, their coefficients became insignificant due to multicollinearity.

First, we have reestimated equation (iv) of Table-1 by adding an index of education. The education index EI is developed by using the primary, secondary and university enrollments. Data on enrollments in other post-secondary institutions are not available. These enrollment ratios to population are weighted by using the Hall and Jones (1999) weights. Details are explained in the appendix. Although current enrollment ratios are not good indicators of the education of workers, this is the best we could do with the available data and therefore EI is unlikely to be a good proxy. We first combined EI and LEI and multiplied employment with the composite index. However, the estimates were unsatisfactory. Furthermore, since our main interest is to estimate the effects of health, a combined education and health index is not useful. Therefore, education index EI is assumed to effect the intercept as in BCS specification and LEI is retained as in equation (iv) of Table-1. Estimates of this augmented equation (vii) are in column (1) of Table-2. It is well determined and compared to equation (iv) in Table-1, showed an improvement. Its $\bar{G}R^2 = 0.542$ is a small improvement on (iv). The most noteworthy feature of (vii) is that the coefficient of profit share increased slightly and the coefficient of the dummies decreased slightly compared to those in (iv). Therefore, the effect of life expectancy implied by this equation at the assumed GDP for 2002 at F\$3,000 millions is F\$38.8 million and this is very close to that implied by equation (iv). Improvement in education, however, seems to have a larger effect on output than health, but its coefficient is insignificant even at the 10% level.

Next, we have reestimated (iv) by augmenting it with a trade openness (TRADE) index and this is measured as the ratio of exports plus imports to the GDP. Estimates are in equation (viii) in column 2 of Table-2. Like equation (vii), this equation showed a small improvement on (iv) with a $\bar{G}R^2 = 0.586$ and the estimates of other coefficients remained the same as in (iv). Only the coefficients of dummies decreased marginally as in (vii). Therefore, there is no significant change in the effects of health on GDP. It is interesting to note that TRADE has a larger effect and a one point increase in trade openness adds an extra \$117 million to GDP. We then added both EI and TRADE to equation (iv). However, neither of their coefficients are significant due to multicollinearity. The estimates are not reported to conserve space.

To minimize multi colleniarity between EI and TRADE, since their correlation coefficient is 0.80, we have computed the principal components with EI and TRADE because our main interest is to estimate the effects of life expectancy on output and the individual effects of EI and TRADE are secondary. Using the first principal component (PC) equation (ix) in Table-2 is estimated. It can be seen that it is well determined and all the χ^2 statistics are satisfactory and $\bar{G}R^2 = 0.625$ is an improvement. The coefficient of PC is close to the coefficient of TRADE in (viii) and significant. In equation (ix) the estimate of profit share slightly increased compared to (iv) implying that the effect of life expectancy on output has only marginally declined to F\$38.9 million. These results thus lend support to the view that the effects of life expectancy on GDP are fairly robust to alternative specifications of the production function.

Finally we have tested for the validity of the BCS constraints viz., that the coefficients of the levels and first differences of the variables are same. These constraints are validated by the Wald test and the

TABLE-2
DETERMINANTS OF OUTPUT IN FIJI
2SNLLS-IV ESTIMATES

Variables	Vii	Viii	ix	x
INTERCEPT	-3.3787 (-26.42)	-3.3235 (-34.42)	-3.3454 (-34.55)	-5.4863 (-4.26)
λ	-1.0976 (-5.36)	-1.1824 (-5.93)	-1.2077 (-6.56)	-1.3329 (-6.76)
EI_{t-1}	0.1274 (1.57)			
$TRADE_{t-1}$		0.0390 (2.54)		
LEI_{t-1}				0.0319 (1.81)
PC_{t-1}			.036747 (2.61)	-0.0638 (-0.66)
$\ln K_{t-1}$	0.2223 (6.26)	0.2076 (6.72)	0.2112 (7.03)	0.1006 (0.84)
$\ln L_{t-1}$				0.8994 (c)
$\ln (L_{t-1} \times LEI_{t-1})$	0.7774 (c)	0.7924 (c)	0.7888 (c)	
$\Delta \ln L_t$				0.8994 (c)
$\Delta \ln K_t$				0.1006
ΔLEI_t				0.0319 (c)
$\Delta \ln(L_t \times LEI_t)$	0.5481 (2.90)	0.6031 (3.25)	0.6000 (3.62)	
$COUP$	-0.0378 (-2.38)	-0.0373 (-2.48)	-0.0371 (-2.56)	-0.0434 (2.40)
OUTLIERS	0.0378 (c)	0.0373 (c)	0.0371 (c)	0.0434 (c)
ΔPC_t				-0.0638 (c)
$\overline{R}^2$	0.561	0.599	0.601	0.472
$\overline{GR}^2$	0.542	0.586	0.625	0.594
Saragan's χ^2	5.501 [0.599]	3.107 [0.875]	3.030 [0.882]	3.969 [0.913]
SEE	0.032	0.031	0.031	0.035
$\chi^2(sc)$	1.867 [0.172]	1.311 [0.252]	1.405 [0.236]	0.815 [0.367]
$\chi^2(ff)$	2.570 [0.109]	0.278 [0.598]	0.390 [0.532]	0.095 [0.758]
$\chi^2(n)$	0.627 [0.731]	1.026 [0.599]	1.062 [0.588]	0.773 [0.679]
$\chi^2(hs)$	2.172 [0.141]	2.630 [0.105]	2.147 [0.143]	1.085 [0.298]

Notes: t-ratios are in the parentheses below the coefficients; p-values are in the square brackets for the χ^2 tests; constrained estimates are denoted with (c).

computed test statistic, with p-value in the square brackets, is $\chi^2(3) =$

2.4285 [0.488]. The equation with the BCS constraints is in (x) of Table-2. It can be seen that this equation is unsatisfactory although its $\bar{G}R^2$ has only decreased slightly. The coefficient of PC is negative and insignificant. The coefficient of capital is low and insignificant. It is likely that these constraints have made the coefficients of the crucial variables insignificant. Therefore, the BCS constraints seem to be somewhat inconsistent with the underlying data generating process (DGP). Needless to say, it is necessary to search for the lag structure in time series data to make the estimates consistent with the underlying DGP and BCS constraints are not useful in the time series estimates.

IV. CONCLUSIONS

In this paper we have developed a framework to estimate the effects of some key determinants of output, besides the factor inputs. We have demonstrated, with time series data from Fiji, how to estimate the effects of life expectancy on output together with two other key variables viz., education and openness of the economy. We have found that policies that increase life expectancy by a year add F$40 million to Fiji's GDP, i.e., about 0.013 percent of the GDP. This is substantially less than 0.04 percent extra growth rate estimate from the BCS cross-country study. Given that Fiji's trend rate of growth of GDP was about 0.0046 (see equation (iii) in Table-1), it is hard to accept that a year increase in life expectancy will increase this trend rate of growth to 0.0446. A simple growth accounting exercise showed that, of the 96.4% increase in output between 1970 and 2002, growth in the two inputs explains 87.2% of this increase, leaving 12.8% as the residual. If this residual is modeled as an increase in TP, the implied annual growth of TP is 0.004 which is approximately the same as the coefficient of trend in equation (iii) of Table-1. Addition of trend, for example, to equation (iv) in Table-1 which allows for the level effects of life expectancy, has significantly reduced the coefficient of trend from 0.0046 in (iii) to 0.0016. This estimate is not reported to conserve space. Therefore, it is unlikely that improvements in life expectancy would have substantial and permanent growth effects.

Our results are robust and did not change significantly when additional determinants of output have been added to the production function. We have also shown that it is preferable to examine the effects of variables that improve human capital through their effects on the measured labour instead of their effects on the intercept technology parameter because of multicollinearity with other variables.

Although our attempt is preliminary and has some limitations, we hope that our framework would be useful to others working with country specific time series data because of the well known limitations of the cross-country studies. However, a main problem with time series data is that many variables are correlated and their individual effects are difficult to estimate. Nevertheless, if the main objective is to estimate the effects of one or two key variables, the rest of the variables can be combined with the principal components procedure. This procedure will go hopefully some way to minimize the omitted variable bias and is likely to give reasonably reliable estimates.

A limitation of our study is that we have not been able to successfully add variables to proxy the effects of institutional and political reforms, other than a dummy for the first political coup of 1987 in Fiji; see footnote 9. Its coefficient is significant and implied that the decrease in output by about 0.035 percent was temporary and lasted from 1987 to 1988. It also indicates that if political stability is restored, it may have a positive effect on output. The major problem in adding other variables is partly due to the difficulty in obtaining reliable data. We hope that other investigators will extend our framework to countries with improved and richer data bases.

Data Appendix

Y is the real gross domestic product in 1990 prices.

L is employment and these are from the Bureau of Statistics (BOS) and Reserve Bank of Fiji (RBF) publications.

K is capital stock, estimated with the perpetual inventory methods with the assumption that the depreciation rate is 4%. The initial capital stock estimate used for 1970 is F$1446.225 million is from Fiji's Economic Development Plan. Investment data used to compute K includes investment in private and public corporate sectors.

$TRADE$ is the ratio of Fiji's total exports of goods and services plus imports of goods services to GDP. The TRADE variable is converted into an index number by assuming that in 1970 it is 1.

LE is life expectancy in years and LEI is the index number of LE with the assumption that in 1970 $LEI = 1$.

EI is an index number which is 1 in 1970. The proportion of enrollments to population of primary, secondary and university enrollments is used to estimate the education levels of the employed workers. Workers with no formal education are given a weight of one. Workers with primary, secondary and tertiary education are given weights of 1.134, 1.244 and 1.312 respectively. The aggregated series is converted into an index number. The weights selected reflect the earnings differences and these are from Hall and Jones (1999). Needless to say this is only an approximation.

PC is the first principal component of $TRADE$ and EI. The loading factors are 0.94173 and 0.83207.

COUP is 1 in 1987, 1988 and 1989. Zero in all other periods.

OUTLIERS is a dummy variable taking a value of 1 in 1995, 1996, 2001 and −1 in 1997. In all other periods it is zero.

Sources of Data:

1. Output, employment and investment data are from the IFS CD-ROM 2003, BOS publications and the RBF Quarterly Review (various issues).
2. Enrollments data are from the Financial Reports for the Ministry of Education (various issues) and from the Planning and Development Office of the USP.
3. Total population data are from Key Statistics, June 2005 issue.
4. Life expectancy data are from the World Bank Indicators, CD Rom, 2004.

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