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The Growth Effects of Education in Australia

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Abstract

The growth effects of human capital, measured in various ways, are controversial and inconclusive. In this paper we estimate the growth effect of human capital with country-specific time series data for Australia. In doing so, we extended the Solow (1956) growth model by using educational attainment as a measure of human capital developed by Barro and Lee (2010). The extended Solow (1956) model performs well after allowing for the presence of structural changes. Our results, based on alternative time series methods, show that educational attainment has a small and significant permanent effect on the growth rate of per worker output in Australia. For comparison of results, alternative measures of human capital are also utilized.

Keywords: SSGR, Economic Growth, Education, Australia.

JEL Classification: C22, O56, O40.

1. Introduction

A very well documented empirical fact is that human capital in its multiple dimensions drives both the creation and application of knowledge and economic growth. Endogenous growth models (ENGMs) have been formulated to investigate whether the variables of interest (for example human capital) yield permanent growth effects. It started in the seminal paper by Romer (1986) who showed that knowledge spillovers have a permanent effect on the growth rate of output. Actually this idea stemmed from Arrow (1962) who argued that externalities arising from 'learning by doing' and knowledge spillovers positively affect labour productivity. Later, Lucas (1988) validated the existing findings that creation of human capital explains total factor productivity (TFP). However, an alternative approach is to extend the Solow's (1956) neoclassical growth model. Using this framework, Mankiw et al. (1992) showed that human capital has permanent level effects. Recently, Rao (2010a) utilized a similar framework to investigate the steady state growth rates (SSGR) for Asian countries.[1]

Following the early work of Barro and his collaborators (Barro, 1991; Barro and Sala-I-Martin, 1995; Barro and Lee, 1996), a large number of growth regressions containing human capital variables in the set of regressors have emerged. These studies employed either cross-section or panel data and can be classified depending on the type of human capital variables they have used. The first group is those that link output growth to some initial level or stock of educational attainment, such as school enrolment rates, for example among others are Barro (1991), Mankiw et al. (1992), Levine and Renelt (1992), Englander and Gurney (1994), Benhabib and Spiegel (1994), Loayza (1994), Caselli et al. (1996) and Hauk and Wacziarg (2004). The second group is those that relate growth to the flow of educational attainment rather than its level, for example Barro and Lee (1993), Graff (1995 & 1996), Barro (1997), Judson (1998), among others. While the first group supports that stock of human capital drives growth, the second group attributes such growth to the accumulation of human capital. Moreover, there are studies that have used altenative measures of human capital based on both stocks and annual average growth rates, for example Gemmell (1996) and De La Fuente and Doménech (2000), among others; they found the latter measure yields plausible estimates.

[1] Rao (2010a) showed that trade openness yields a permanent effect on the growth rate of output in the Asian countries.

The time series evidence on the impact of human capital on growth is inadequate, perhaps due to unavailability of consistent data on education and training variables. The recent attempts that used time series data include Jenkins (1995), Asteriou and Agiomirgianakis (2001), Rao and Vadlamanati (2010) and Leoning et al. (2010). In the case of the UK, Jenkins (1995) found that highly qualified workers contribute almost twice as much to productive efficiency as those with no qualifications. Three proxies for the stock of human capital were developed via considering workforce qualifications. Asteriou and Agiomirgianakis (2001) attained a statistically significant relationship between primary, secondary and higher education enrolments and GDP per capita for Greece. Rao and Vadlamannati (2010) showed that human capital (measured as secondary school enrolment ratio) has both a permanent level and a permanent growth effect in India. Using data from Guatemala, Leoning et al. (2010) found that human capital (measured as average year of total schooling) has a highly significant and positive impact on growth. For a comprehensive review on human capital and growth, see Descy and Tessaring (2004).

In this paper, we contribute to this literature on three different fronts. First, we apply alternative time series techniques to estimate the SSGR for Australia over the past 50 years, with a particular focus on the contribution of human capital on growth. This is of special interest because there are only a few studies that have estimated and analyzed the SSGR using country-specific time series data. Empirical works on growth are mostly based on cross-country analysis and to this end country-specific time series studies are more appealing since they overcome the heterogeneity problem and take into account the unique historical information for each country.[2] Second, it is noted that the measurement of human capital in most empirical works is not satisfactory; a frequently used measure is the enrolment rates in primary, secondary or tertiary education.[3] According to Bergheim (2008), enrolment rate is not a useful measure of human capital because it does not include information on years of education.[4] We show that alternative measures (total school enrolment rate, average year of

[2] The cross-section study may also be inadequate if returns to education differ substantially across countries.
[3] Secondary (primary or tertiary) enrolment is the percentage of the number of people undertaking secondary (primary or tertiary) education in a given year with respect to the total number of people present in the age group.
[4] For example let's assume that two countries (for instance A and B) have same secondary enrolment rates (about 70%) but different stock of human capital (years of education). If country A has lower stock (5 years of education), 70% secondary enrolment rate will lead to a huge rise in the average years of education in the workforce. On the other hand, if country B has high stock (12 years of education), 70% secondary enrolment rate may not be sufficient to maintain the initial level of human capital. To this end, we need information about the

primary schooling, average year of secondary schooling and average year of tertiary schooling) understates the growth effect of human capital in Australia. To this end, average year of total schooling (educational attainment) seem to yield plausible results. Finally, it is imperative to consider structural changes when estimating the level and growth effects of human capital. For the purpose of robustness in the results, structural changes must be addressed.

The structure of this paper is as follows. Section 2 discusses our extensions to the Solow model and develops our specifications. Empirical results are discussed and presented in Section 3. Finally, Section 4 concludes.

2. Specification

The starting point is the steady state solution for the level of output in the Solow (1956) growth model and this is:

$$y^* = \left(\frac{s}{d+g+n}\right)^{\frac{\alpha}{1-\alpha}} A \qquad (1)$$

where y^* = steady state level of income per worker, s = ratio of investment to income, d = depreciation rate of capital, g = rate of technical progress, n = rate of growth of labour, A = stock of knowledge and α = exponent of capital in the Cobb-Douglas production function with constant returns. This implies that SSGR, assuming that all other ratios and parameters are constant, is simply TFP because:

$$\Delta \ln y^* = SSGR = \Delta \ln A = TFP \qquad (2)$$

Since the determinants of TFP are not known and are exogenous in the Solow (1956) growth model, it is therefore the Solow model is also known as the exogenous growth model (EXGM). The new growth theories based on the ENGMs use optimization framework and

initial stock and combine the two measures to get a sense for the future path of human capital, for example the average years of education of the working age population.

suggest several potential determinants of TFP. However, to the best of our knowledge it is hard to argue, as asserted by Parente (2001), that ENGMs are empirically better than the extended EXGMs. We take the view that Solow model can be extended by making TFP a function of the potential determinants identified by the ENGMs. Furthermore, the extended Solow model is much simpler to estimate, on the other hand, it is necessary to use a complex system of non-linear dynamic equations to estimate a standard ENGM; see Greiner et al. (2005) for more details on ENGMs.[5] We are not aware of any ENGM in which the functional form of the determinants of TFP is well established with theoretical insights.

We extend the Solow model to estimate the SSGR as follows. Note that the SSGR can be estimated by estimating an extended production function by assuming that the stock of knowledge (A) depends on some important variables identified by the ENGMs. We start with the well-known Cobb-Douglas production function (Y = output, K = capital stock and L = labour) with constant returns:

$$Y_t = A_t K_t^{\alpha} L_t^{(1-\alpha)} \qquad (3)$$

Generally in empirical works A is assumed to evolve as $A_t = A_0 e^{gt}$ where A_0 = initial stock of knowledge and g = growth rate of A per period and t = time. Following Rao (2010b) and Paradiso and Rao (2011) we can modify this evolution in two ways by making g a simple or a non-linear function of HKI (human capital measured as average year of total schooling) as follows.

$$g = (\varpi HKI)t \qquad (4)$$

$$g = (\varpi HKI)t + \gamma_1 HKI + \gamma_2 HKI^2 \qquad (5)$$

[5] In the seminal contribution by Barro (1991), his growth equation is mistaken to be based on some unknown ENGM. Actually Barro's growth equation is based on the human capital augmented version of the Solow model (see Mankiw et al., 1992) and the adjustment equation they propose makes growth of output to adjust to the gap between steady state and actual output i.e., $\Delta \log y = \lambda(\log y^{*} - \log y)$, where y = actual output. Therefore, Barro's growth equation is useful to measure the growth rate during the transition period, however it is less helpful in deriving the permanent growth effects of the variables of interest because transitional growth vanishes in the steady state.

These formulations for g are based on empirical considerations and in our case specification (5) gave the best empirical results. In equation (4) ϖ measures SSGR due to HKI. The SSGR effects of HKI are assumed to have some dynamic component in equation (5), which are captured by HKI and HKI^2. Substitution of (5) into (3) in its intensive form gives:

$$y_t = A_0 e^{\left(\varpi HKI_t \cdot T + \gamma_1 HKI_t + \gamma_2 HKI_t^2\right)} k_t^{\alpha} \tag{6}$$

where $y = (Y/L)$ and $k = (K/L)$. Expressing the evolution of the stock of knowledge A as modified in (5) in log terms and denoting logs with lower case letters, we have:

$$a_t = a_0 + \varpi HKI_t \cdot T + \gamma_1 HKI_t + \gamma_2 HKI_t^2 \tag{7}$$

Taking the first difference gives:

$$\begin{aligned} \Delta a_t = TFP = \varpi_1 \Delta HKI_t \times T + \varpi_2 HKI_t \\ + \gamma_1 \Delta HKI_t + \gamma_2 \Delta HKI_t \times HKI_t \end{aligned} \tag{8}$$

Equation (8) can be interpreted as the intermediate period effects of *HKI* on SSGR.[6] In the long run, however, all the differences of the variables become zero in the steady state. Therefore, the SSGR is:

$$\therefore SSGR = \varpi HKI \tag{9}$$

Based on equation (9), it could be asserted that the higher is *HKI*, the higher becomes the SSGR.

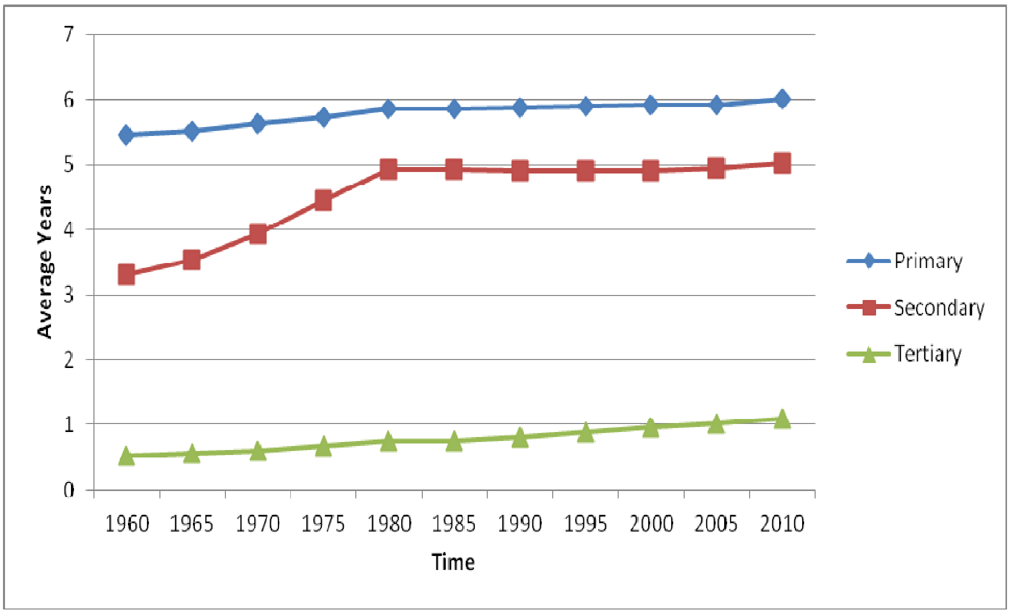
[6] In this formulation, both the level and change of HKI have growth effects and this is consistent with the growth accounting approach of Benhabib and Spiegel (1994). However, the specifications are not identical.

3. Empirical Results

Some statistical considerations

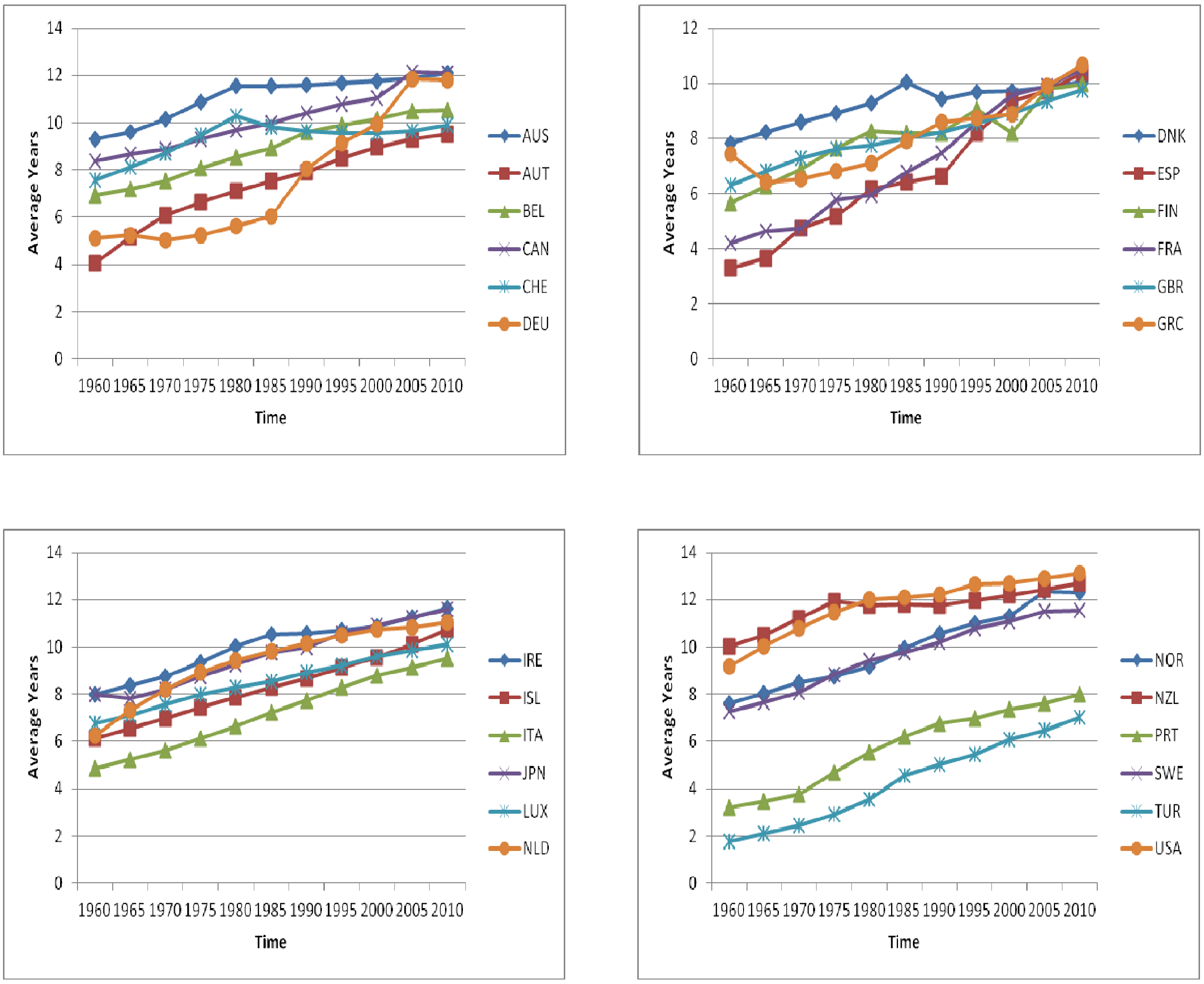
We now briefly discuss the broad trends in the variables of interest to provide a backdrop and to discuss the policy implications of our findings. During 1960-2008, Australia has experienced an average GDP growth rate of 3.5%. Although during this period Australia had encountered significant structural changes (for example, among others were three recessions of varied scale (1974, 1982 and 1990-1991) and a monetary policy regime shift in 1996 with the introduction of inflation targeting[7]), its growth rate has remained well above 2.5% per year. The average growth rates over the sub-periods 1960-69, 1970-79, 1980-89, 1990-99 and 2000-08 are 5.5%, 3.1%, 3.4%, 3.3% and 3.1%, respectively. These growth rates are reasonable and attained partly due to the reform policies, detainment of strong social services and improvements in education and training.

Figure 1: Average years of primary, secondary and tertiary education in Australia



[7] In Australia inflation targeting was first adopted by the Reserve Bank of Australia in 1993, as an operational interpretation of the price stability goal of its legislated mandate. The inflation targeting framework was subsequently verbally endorsed by the government of the day, but was not formally endorsed until 1996, when a new government signed a letter of agreement with the new Governor.

Figure 2: Evolution of the average years of total schooling in advanced countries



NB: AUS = Australia; AUT = Austria; BEL = Belgium; CAN = Canada; CHE = Switzerland; DEU = Germany; DNK = Denmark; ESP = Spain; FIN = Finland; FRA = France; GBR = Great Britain; GRC = Greece; IRE = Ireland; ISL = Iceland; ITA = Italy; JPN = Japan; LUX = Luxembourg; NLD = Netherlands; NOR = Norway; NZL = New Zealand; PRT = Portugal; SWE = Sweden; TUR = Turkey and USA = United States of America.

Economic reforms in Australia are always complemented by policies to provide the skills and training needed in the technologically-sophisticated economy, for instance, technical advancements in the banking sector created considerable opportunities for on-the-job and off-the-job training. Since the 1980s, retention rates in the secondary education dramatically increased followed by a sharp increase in enrolments in vocational colleges and universities. By 2002, education expenditure as a proportion of GDP had caught up with the average of member countries of the OECD; Australia 6%, OECD 5.8% and USA 7.2% (OECD, 2005). According to the Barro and Lee (2010) dataset, average educational attainment in Australia is 12.12 years in 2010. Figure 1 illustrates the average attainment with respect to primary, secondary and tertiary education.[8] From 1960 to 2010, the average year of primary schooling is the highest up to 6 years, while the average year of secondary schooling is between 3 to 5 years. The average attainment in tertiary education has been the lowest and since 2000 it has reached 1 year. Figure 2 illustrates the evolution of HKI[9] in the last 50 years for advanced countries including Australia. It shows that all countries have different but close HKI particularly since the 1990s. The only exceptions are Portugal and Turkey which has much lower *HKI* than other countries. The Australian *HKI* is consistent with schooling levels in other leading countries such as New Zealand and the USA. To this end, educational attainment could have played an important role in explaining the long-term growth rate or the SSGR of Australia. We investigate this aspect with an extended version of the Solow (1956) growth model.

Unit root tests

Lee and Strazicich's (2003) two break minimum *LM* unit root tests were applied to assess the order of integration of the variables. The break dates are endogenously determined and can be explained using two models i.e., model A and model C. These models are based on alternative assumptions about structural breaks, for instance model A allows for two shifts in the intercept and model C includes two shifts in the intercept and trend. Table 1 displays the results of these tests. The test statistics of the *LM* unit root tests for the three variables (y, k and *HKI* in levels) do not exceed the critical values in absolute terms and therefore the unit root null cannot be rejected at the 5% level. For the first differences of these variables the unit root null is rejected at the 5% level. The t-statistics corresponding to the break dates are statistically significant at the conventional levels (not reported for brevity).

In most cases the break dates are located during the 1980s and 1990s. These are consistent with the timings of macroeconomic events that was experienced by the Australian economy, for instance, large per capital income fluctuations (1970s), recessions (early 1970s, 1980s and 1990s), education reform policies especially on adult literacy (1996), deregulation policies

[8] Barro and Lee (2010) data is used to construct Figures 1 and 2.
[9] The dataset of Barro and Lee (2010) has observations every 5 years between 1950 and 2010. Intermediate data are linearly interpolated, for example see Bergheim (2008), Park (2010) and Besley and Reynal-Querol (2011). Since the evolution of this variable over time is quite stable, simple linear interpolation to construct annual data does not create problems or distortions.

and the Australian dollar float (mid 1980s), formation of the Australian Stock Exchange Limited (1987), and greater openness and microeconomic reforms (since 1990s).

Table 1: Two-break minimum *LM* unit root test, 1960-2008

Variables	Level				First Difference			
	Model A		Model C		Model A		Model C	
	Test statistic	Break dates	Test statistic	Break dates	Test statistic	Break dates	Test statistic	Break dates
y	-0.415 [4]	1981; 1995	-0.869 [4]	1975; 1995	-4.963 [3]	1991; 2004	-5.307 [5]	1985; 1992
k	-1.113 [3]	1987; 1991	-0.182 [5]	1974; 1992	-5.002 [3]	1975; 1995	-5.376 [3]	1995; 2003
HKI	-0.941 [6]	1976; 1989	-1.601 [4]	1996; 1980	-4.128 [5]	1975; 1986	-6.372 [4]	1980; 1996

NB: The 5% critical values for models A and C are -3.842 and -5.286, respectively. The number in square brackets indicates the optimal number of lagged first-differenced terms included in the unit root test to correct for serial correlation. Critical values are taken from Lee and Strazicich (2004, 2003). Kumar et al. (2013) contain more details on this test. RATS 7.2 was used to used to perform this test.

Estimates without structural changes

The unit root tests strongly indicates that the series are I(1) in levels, it is therefore necessary to estimate the extended Solow (1956) growth model using time series cointegration techniques. We utilized four techniques viz. canonical cointegrating regression (CCR), general to specific (GETS), dynamic ordinary least squares (DOLS) and fully modified ordinary least squares (FMOLS). These techniques are classified as single-equation estimators and they deal with the problem of second-order asymptotic bias arising from serial correlation and endogeneity and are asymptotically equivalent and efficient. Park (1992) proposed the CCR technique which is simple to apply, and as efficient as methods based on system maximum likelihood estimation. The CCR technique is quite similar to Phillip and Hansen's (1990) FMOLS. While the former selects a canonical regression among the class of models representing the same cointegrating relationship, the latter modifies variables and estimates directly to eliminate the existing nuisance parameters. Operationally, the CCR method concentrates on the data transformations, but FMOLS use the transformations of both the data and estimates. In contrast, Stock and Watson's (1993) DOLS method is parametric and a form of distributed lag approach that involves the inclusion of lags and leads of the first differences of the explanatory variables as part of the regressors. The GETS technique was proposed by the London School of Economics Professor David Hendry and it utilizes the general dynamic specification similar to the autoregressive distributed lag model. The variable deletion tests are applied to attain the parsimonious estimated model; for more details on the GETS technique, see Rao et al. (2010).

The general form of the extended Solow model is given as:

$$\ln y_t = Intercept + \alpha \ln k_t + \gamma_1 HKI_t + \gamma_2 HKI_t^2 + \varpi HKI \cdot T \qquad (10)$$

In the first instance we estimated equation (10) without allowing for any structural changes that was experienced in the domestic economy. Table 2 presents these results.

Table 2: FMOLS, CCR, DOLS and GETS estimates without dummies, 1960-2008

$\ln y_t = Intercept + \alpha \ln k_t + \gamma_1 HKI_t + \gamma_2 HKI_t^2 + \varpi HKI \cdot T$	FMOLS	CCR	DOLS	GETS
Intercept	-10.696	-10.481	-9.117	-3.898
	[5.125]***	[5.808]***	[2.779]**	[2.337]**
α	0.323	0.298	0.673	0.391
	[0.947]	[0.805]	[1.834]*	[0.864]
γ_1	1.580	1.528	1.465	1.461
	[3.122]***	[3.259]***	[2.097]**	[2.354]**
γ_2	-0.075	-0.073	-0.068	-0.070
	[3.056]***	[3.175]***	[1.919]*	[2.332]**
ϖ	0.001	0.001	0.001	0.001
	[2.991]***	[2.828]**	[1.849]*	[2.066]**
λ		-0.368		-0.397
		[4.529]***		[3.121]***
EG residual test		-3.784***		-
LM(1) test (p-value)		0.304		0.635
LM(2) test (p-value)		0.511		0.380
LM(4) test (p-value)		0.632		0.702
JB test (p-value)		0.450		0.737
BPG test (p-value)		0.107		0.386

NB: The *t*-statistics are in [ ] brackets. *, **, *** denote significance at 10%, 5%, and 1%, respectively. FMOLS = fully modified ordinary least squares; CCR = canonical cointegrating regression; DOLS = dynamic ordinary least squares; GETS = general to specific; and EG = Engle-Granger *t*-test for cointegration. λ , factor loading in the ECM. BPG = Breusch-Pagan-Godfrey heteroskedasticiy test; JB = Jarque-Bera normality test; LM = Bresuch-Godfrey serial correlation LM test. FMOLS uses Newey-West automatic bandwidth selection in computing the long-run variance matrix. In the DOLS leads and lags are selected using the AIC criteria. The standard errors (not reported) for the DOLS estimation are calculated using the Newey-West correction. The GETS equation was estimated using non-linear least squares as follows: (r squared was 0.41 and due to short sample only one lag was used):

$$\Delta \ln y_t = Intercept + \sum_{i=1}^{n1} \mu_{1i} \Delta \ln y_{t-i} + \sum_{i=1}^{n2} \mu_{2i} \Delta \ln k_{t-i} + \sum_{i=1}^{n3} \mu_{3i} \Delta HKI_{t-i}$$
$$+\lambda \left[\ln y_{t-1} - \left(Intercept + \alpha \ln k_t + \gamma_1 HKI_t + \gamma_2 HKI_t^2 + \varpi HKI \cdot T \right) \right]$$

All tests were performed using Eviews 7.0 software.

Fairly consistent estimates were attained across the four estimators. The speed of adjustment (λ) implies negative feedback mechanism and is statistically significant at 1% level. The Engle and Granger (1987) t-test supports the existence of cointegration among the variables at 1% level. Moreover, the diagnostic tests indicate no issues with respect to serial correlation, normality and heteroscedasticity. The growth effect of *HKI* is 0.001 and statistically significant at the conventional levels. In GETS, CCR and FMOLS the capital share is between 0.3 to 0.4, however the DOLS technique produced implausibly high estimate at around 0.7. Further the estimates of capital share are statistically insignificant at conventional levels in all cases, except in DOLS at 10% level. While the results suggest that human capital has permanent growth effects, it is difficult to assert that the findings are robust because the capital-output ratios are statistically insignificant at the conventional levels.[10] To achieve robust estimates, we tested for structural changes and introduced various dummy variables in the extended Solow (1956) model.

Structural change tests

We tested for stability of the estimated equations in Table 2. In doing so, we applied the Quandt (1960) and Andrews (1993) structural break tests. In the time series context, techniques such as Gregory and Hansen (1996a & b) and Arai and Kurozumi (2007) are well suited to test for stability of an estimated relationship, however they are difficult to implement in growth models particularly in growth effect estimations where trend is included in the specification. Further it is also difficult to estimate the extended Solow (1956) model with regime shifts. To this end, we employed the simple test proposed by Quandt-Andrews to identify the significant structural breaks in our estimated equations. Since this test performs only when the parameters are linear, we utilize the OLS estimates of GETS for this purpose.[11]

Prior to further discussion, it would be useful to take an overview of the Quandt-Andrews test. Based on Quandt (1960), Andrews (1993) modified the Chow test and allows for unknown breakpoints in the sample for an estimated equation. It utilizes the Chow breakpoint tests and this is performed at every observation over the interval $[\xi T, (1-\xi)T]$ and calculates the

[10] Except the DOLS estimate at 10% level.
[11] These estimates are not significantly different from the estimates reported in Table 3. We did not report these estimates but can be obtained from the authors upon request.

supremum of the F_k statistics as $\sup F = \sup_{k \in [\xi T, (1-\xi)T]} F_k$ where ξ is a trimming parameter. Andrews and Ploberger (1994) developed two additional test statistics i.e. the average (ave F) and the exponential (exp F). The null hypothesis of no break is rejected if these test statistics are large, however Hansen (1997) derives an algorithm to compute approximate asymptotic p-values of these tests.

Table 3: Quandt-Andrews structural break tests, 1960-2008

Statistic	Value	Break Date	Prob.
Maximum LR F-statistic	3.759	1996	0.002***
Maximum Wald F-statistic	23.348	1974	0.000***
Exp LR F-statistic	7.179	-	0.000***
Exp Wald F-statistic	18.208	-	0.000***
Ave LR F-statistic	5.032	-	0.001***
Ave Wald F-statistic	23.106	-	0.000***

NB: Probabilities calculated using Hansen's (1997) method. *** indicates significance at 1% level. Eviews 7.0 was used to perform this test.

Table 4: Chow structural break tests, 1960-2008

Event	Break Date	Test Statistics		
		F-statistic	LL ratio	Wald statistic
Peak in manufacturing sector	1965	0.452 (0.83)	3.492 (0.74)	-
Oil price shocks	1973	1.141 (0.35)	8.358 (0.21)	6.581 (0.36)
Surge in wages	1974	3.643 (0.01)	19.745 (0.00)	12.368 (0.05)
Recession	1982	1.223 (0.31)	8.908 (0.17)	66.187 (0.00)
Financial deregulation and Australian dollar float	1985	2.092 (0.04)	14.360 (0.02)	12.644 (0.05)
Formation of Australian Stock Exchange Limited	1987	1.916 (0.10)	8.083 (0.23)	10.177 (0.11)
Recession	1990	2.619 (0.03)	17.391 (0.01)	15.455 (0.02)
Asian financial crises	1997	3.105 (0.01)	20.023 (0.00)	12.386 (0.05)
Introduction of goods and services tax	2000	1.717 (0.14)	7.995 (0.17)	12.377 (0.05)
Language, literacy and numeracy programme	2002	0.865 (0.52)	6.466 (0.37)	4.843 (0.56)

NB: LL means log likelihood ratio. Probability values are in parentheses. – indicates not available due to short sample. Eviews 7.0 was used to perform this test.

The Quandt-Andrews test results are reported in Table 3. All test statistics (maximum, exponential and average) reject the null of no structural breaks at 1% level. The detected break dates are 1974 and 1996 and these are not unrealistic because Australia experienced a recession during 1974 and 1996 signifies the introduction of inflation targeting regime in the conduct of monetary policy. Moreover, there are a number of other structural changes that took place in Australia and it is vital to account for these shifts in the growth model. To test the significance of these additional structural changes, we employ Chow's (1960) exogenous breakpoint tests. If the potential breakpoint is known a priori, it is suitable to use this method to test the null of no structural break against the alternative of a break at that time. Table 4 present the results of Chow's breakpoint test associated with some key structural changes in

Australia. All three test statistics reject the null of no breaks at specified breakpoint for the following cases: i. 1985 financial deregulation and Australian dollar float; ii. 1990 recession; iii. 1974 surge in wages[12]; and iv. 1997 Asian financial crisis. For 1982 recession and 2000 introduction to goods and services tax only Wald statistic rejected the null of no break at 5% level. Further, F-statistic rejected the null of no break at 10% level for 1987 formation of Australian Stock Exchange Limited. Consequently, these structural changes are modelled as dummy variable regressors in the extended Solow model.

Estimates with structural changes

The presence of structural changes has led us to estimate the extended Solow (1956) model by including relevant dummy variables. Initially, we included all dummy variables as regressors i.e. 1974 and 1996 from Quandt-Andrews test and 1982, 1985, 1987, 1990, 1997 and 2000 from Chow test, however only three dummies (1974 (DUM74), 1990 (DUM90) and 1996 (shift96)) were statistically significant at the conventional levels and seemed to improve the overall results.[13] This implies that introduction of inflation targeting regime (in 1996) and the two recessions (in 1974 and 1990) had positive and negative impacts on output growth, respectively. The results of the extended Solow model with these dummies are reported below in Table 5.

Application of FMOLS, CCR, DOLS and GETS produced estimates that are plausible and statistically significant at the conventional levels. Note that introducing the dummies altered the magnitude of the estimates only marginally, except the capital share now range between 0.32 to 0.48. Interestingly, the estimates of capital share have become statistically significant and the adjustment coefficient has increased to around -0.8. The Engle-Granger t-test confirms the existence of cointegration among the variables. There are also no issues in terms of diagnostic tests, except for heteroscedasticity in the CCR model but it is not significant at 5% level. In Figure 3 we present the actual and fitted values of $\Delta \ln y$ and the fit is satisfactory implying that the estimates are robust.

[12] There is high probability that this could be capturing recession that hit Australia in 1974.
[13] The results with all dummies are not reported to conserve space but can be obtained from the authors upon request.

Table 5: FMOLS, CCR, DOLS and GETS estimates with dummies, 1960-2008

$$\ln y_t = Intercept + \alpha \ln k_t + \gamma_1 HKI_t + \gamma_2 HKI_t^2 + \varpi HKI \cdot T + \varphi_1 Shift96 + \varphi_2 DUM74 + \varphi_3 DUM90$$

	FMOLS	CCR	DOLS	GETS
Intercept	-11.301	-11.056	-10.937	-9.139
	[13.291]***	[13.303]***	[6.203]***	[5.282]***
α	0.315	0.339	0.429	0.475
	[2.384]**	[2.031]**	[3.035]***	[2.154]**
γ_1	1.663	1.624	1.715	1.801
	[8.165]***	[7.502]***	[5.362]***	[5.909]***
γ_2	-0.078	-0.076	-0.101	-0.085
	[7.860]***	[7.218]***	[4.930]***	[5.751]***
ϖ	0.001	0.001	0.001	0.001
	[5.260]***	[3.735]***	[3.495]***	[2.318]**
Shift96	0.054	0.059	0.057	0.045
	[5.881]***	[5.125]***	[4.374]***	[3.429]***
DUM74	-0.025	-0.023	-0.029	-0.027
	[3.278]***	[2.622]**	[3.116]***	[2.896]**
DUM90	-0.034	-0.032	-0.034	-0.028
	[3.518]***	[2.660]**	[4.997]***	[2.217]**
λ		-0.766		-0.789
		[4.128]***		[6.409]***
EG residual test		-5.031***		-
LM(1) test (p-value)		0.794		0.806
LM(2) test (p-value)		0.864		0.879
LM(4) test (p-value)		0.989		0.743
JB test (p-value)		0.499		0.565
BPG test (p-value)		0.062**		0.980

NB: The *t*-statistics are in [ ] brackets. *, **, *** denote significance at 10%, 5%, and 1%, respectively. FMOLS = fully modified ordinary least squares; CCR = canonical cointegrating regression; DOLS = dynamic ordinary least squares; GETS = general to specific; and EG = Engle-Granger *t*-test for cointegration. λ, factor loading in the ECM. BPG = Breusch-Pagan-Godfrey heteroskedasticiy test; JB = Jarque-Bera normality test; LM = Bresuch-Godfrey serial correlation LM test. FMOLS uses Newey-West automatic bandwidth selection in computing the long-run variance matrix. In the DOLS leads and lags are selected using the AIC criteria. The standard errors (not reported) for the DOLS estimation are calculated using the Newey-West correction. The GETS equation was estimated using non-linear least squares as follows: (r squared was 0.46 and due to short sample only one lag was used):

$$\Delta \ln y_t = Intercept + \sum_{i=1}^{n1} \mu_{1i} \Delta \ln y_{t-i} + \sum_{i=1}^{n2} \mu_{2i} \Delta \ln k_{t-i} + \sum_{i=1}^{n3} \mu_{3i} \Delta HKI_{t-i} + \varphi_1 Shift96 + \varphi_2 DUM74 + \varphi_3 DUM90$$

$$+\lambda \left[\ln y_{t-1} - \left(Intercept + \alpha \ln k_t + \gamma_1 HKI_t + \gamma_2 HKI_t^2 + \varpi HKI \cdot T \right) \right]$$

All tests were performed using Eviews 7.0 software.

Since all techniques yield consistent results, we are confident that our model is correctly specified. The estimate of growth effects of *HKI* is 0.001 and hence we use this value to compute the dynamics of SSGR (see equation 9). The plot of SSGR and the actual growth of output per worker (DLYL) for the last 30 years are presented in Figure 4. The average value of SSGR is around 1% over the period 1960 to 2008. More importantly, this result is in line with a value of 0.96% we found if we use data from Maddison (1995) to calculate an historical average TFP growth rate for Australia for the period 1950 to 1995, and these are consistent with Ferreira et al. (2005).[14] These studies have used the growth accounting procedure to derive their findings.

Figure 3: Actual and fitted series of Δln*y*

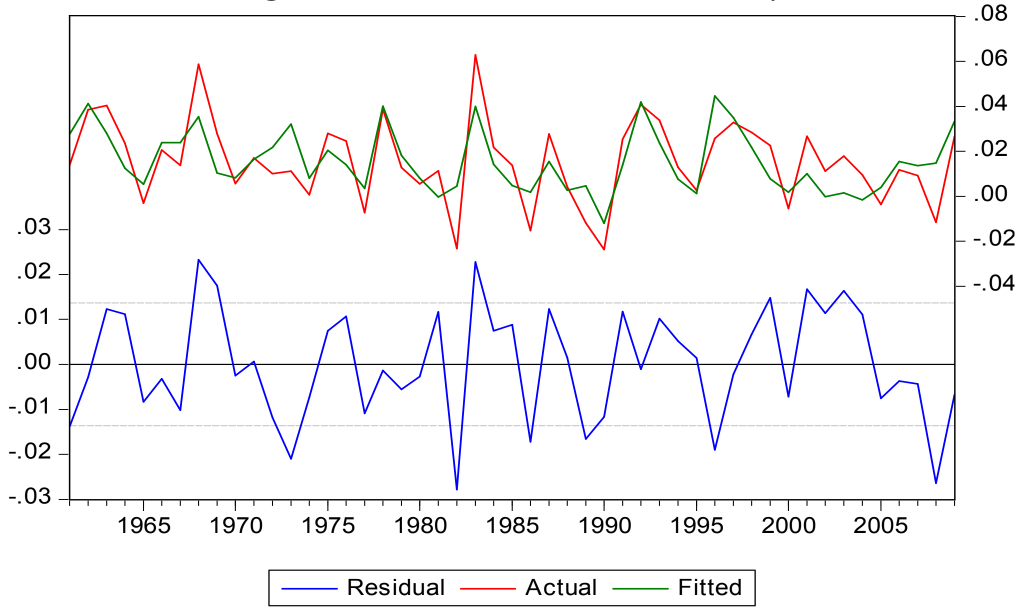
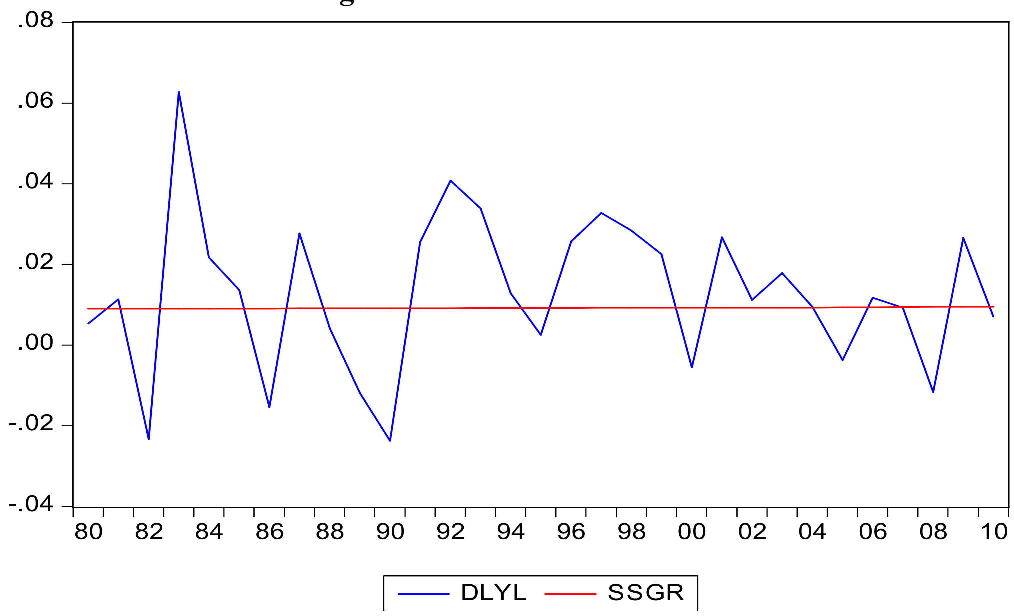


Figure 4: SSGR for Australia



Note: $SSGR = 0.001 * HKI$

[14] Compared to the estimates of SSGR for the developing countries estimates of SSGR for the advanced countries seem to be limited.

Estimates with alternative measures

Aside from using the average year of total schooling as a measure of human capital, we also utilized alternative measures such as the total school enrolment rate, average year of primary schooling, average year of secondary schooling and average year of tertiary schooling to determine the SSGR. To conserve space, we do not tabulate these results but briefly discuss here.[15] The data on total school enrolment rate is retrieved from the World Development Indicators. Due to short sample (1971-2008), we used only GETS and CCR techniques and both yield consistent results. Two dummies (1996 inflation targeting regime and 1990 recession) were incorporated into the model and both were statistically significant at the 5% level. The magnitude of growth effect of total school enrolment rate is 0.00038 and the average value of SSGR is around 0.3% which seems very trivial. The capital share is around 0.4 and statistically significant at the 5% level.

Given that the SSGR due to average year of total schooling is around 1%, it is vital to investigate how much of this is attributed to attainments in primary, secondary and tertiary education.[16] Using four estimation techniques (GETS, CCR, FMOLS and DOLS), we have estimated the individual growth effects of these three variables. The dummy variables used were 1974 recession (only in growth model for average year of primary schooling), 1985 financial deregulation, 1990 recession and 1996 monetary policy shift in all cases. The capital share is between 0.25 to 0.41 and statistically significant at the conventional levels. The estimate of growth effect of average year of primary (tertiary) schooling is 0.00057 (0.00042) and to this end the SSGR is around 0.5% (0.4%). In contrast, the estimate of growth effect of average year of secondary schooling is 0.00030. With regard to the SSGR due to average year of secondary schooling, it is very low at around 0.1%. In all cases, the estimates of growth effects of *HKI* based on the three measures are statistically significant at the 5% level. From a comparative perspective, we argue that average year of total schooling is the optimal measure of human capital and yields relatively higher value of SSGR.

[15] These results can be obtained from the authors upon request.

[16] Note that Barro and Lee's (2010) data on educational attainment is the sum of average attainments in primary, secondary and tertiary education. This led us to estimate 12 equations i.e. three extended versions of Solow (1956) model (with average year of primary schooling, average year of secondary schooling and average year of tertiary schooling) are estimated using four methods.

4. Conclusion and Policy Implications

This paper used an extended Solow (1956) growth model to estimate the long run growth rate for Australia for the period 1960-2008. The endogenous two break minimum *LM* unit root tests revealed that the level variables are non-stationary and provided break dates that are located mostly during the 1980s and 1990s. Four time series techniques (CCR, GETS, FMOLS and DOLS) were utilised to estimate the cointegrating equations. First, we estimated the cointegrating equations without allowing for structural changes. We attained less robust results; capital share was implausibly high in DOLS (around 0.7) and statistically insignificant at the conventional levels in all cases.[17] Second, we employed the Quandt-Andrews and Chow breakpoint tests to investigate the breakpoints in the cointegrating equations. The Quandt-Andrews test rejected the null of no breakpoints and indicated two breakpoints i.e. 1974 (recession) and 1996 (monetary policy shift). Since Chow method tests for exogenous breakpoints, we tested for a number of expected breaks. To this end, several breakpoints were not rejected i.e. 1974 (surge in wages or recession), 1982 (recession), 1985 (financial deregulation and Australian dollar float), 1987 (formation of Australian Stock Exchange Limited), 1990 (recession), 1997 (Asian financial crisis) and 2000 (introduction of goods and services tax).

Third, we estimated the cointegrating equations considering the presence of structural changes depicted by Quandt-Andrews and Chow tests. These structural changes were introduced into the extended Solow model as dummy variable regressors. However, we found that only three dummies viz. 1974 and 1990 recessions and 1996 monetary policy shift were statistically significant at the conventional levels. Further, allowing for these structural changes in the extended Solow model has led us to achieve robust estimates across the four techniques (FMOLS, CCR, DOLS and GETS). The capital share is from 0.32 to 0.48 and has become statistically significant. More importantly, the estimate of growth effect of *HKI* is 0.001 and the average value of SSGR is around 1% over the period 1960 to 2008; this is comparable to Maddison (1995) and Ferreira et al. (2005) where growth accounting procedure was employed to derive the results.

[17] Except the DOLS estimate which is significant at the 10% level.

Lastly, aside from using the average year of total schooling as a measure of human capital, we also utilized alternative measures such as the total school enrolment rate, average year of primary schooling, average year of secondary schooling and average year of tertiary schooling to determine the SSGR. We found that total school enrolment rate understates the SSGR (about 0.3%). Further, the SSGR due to average year of primary schooling and average year of tertiary schooling are around 0.5% and 0.4%, respectively. With respect to the SSGR due to average year of secondary schooling, it is very low at around 0.1%. This implies that the SSGR of 1% (due to average year of total schooling) is mainly attributed to attainments in primary and tertiary education. In light of these findings, we argue that average year of total schooling is the optimal measure of human capital and yields relatively higher value of SSGR for Australia.

The fact that human capital measured as average year of total schooling has a permanent growth effect in Australia implies that meaningful advice for policy makers can be drawn. To increase the SSGR via improving human capital raises an important question: how educational attainment can be increased in Australia? While the average years of primary and tertiary schooling has contributed to SSGR in a somewhat satisfactory way, the contribution of average year of secondary schooling is very trivial. It is well known that reforms are vital to improve attainment rate in secondary education. Policy makers should establish systematic student counselling and career guidance services to prevent a lack of awareness of future options, and in all upper secondary schools to assist students to overcome their problems and prevent dropout. Further, it is important to prepare students well for the transition from basic to upper secondary school to enhance their successful rate. The Council of Australian Government's (COAG's) target to lift the Year 12 or equivalent attainment rate to 90% by 2020 seems reasonable. Other policy directions (for instance, improving teacher and school leader quality, high standards and expectations, greater accountability and better directed resources, modern world class teaching and learning environments including ICT, integrated strategies for low socio-economic status school communities and boosting parental agreement) proposed by COAG will also promote educational attainment in the medium to long-term.

Moreover, there is also an aspiration for a sustainable increase in public spending in education and training sectors. In the 2011-2012 budget, government announced to make 'every school a great school' and has allocated \$425m to reward top performing teachers, \$558m to deliver tailored, quality training places through the National Workforce Development Fund and ambitious reform of vocational education and training, with \$1.75b on offer to partner with the states and territories. While these current efforts are desirable, there could be much more delivered to encourage young people to stay in the education system, especially secondary and tertiary; for example, more scholarships to complete university education, creation of new vocational colleges, greater resources and funding for rural schools etc. Further, it is imperative to frequently assess educational policy outcomes, for instance, whether the implemented reforms have been effective. Although COAG's human capital agenda presents a framework to assess educational reforms, there is a greater need to focus on educational attainment.

Needless to say, there are limitations in this paper. Firstly, although the average year of total schooling as a measure of human capital (retrieved from Barro and Lee, 2010) gave plausible results, it does not include accumulation of knowledge and skills that are attributed to on-the-job training and community-based workshops. Further, human capital also seems to be affected by life expectancy and health care provisions. To this end, there is need to develop a more comprehensive measure of human capital that takes into account all relevant data. Secondly, we did not consider the structural regime shifts in the extended Solow (1956) model. However, we are not aware of any structural break technique that could suitably test for regime shifts in this model. Perhaps this is due to the existence of non-linear parameters and also the way growth effect parameter is formulated. A pragmatic approach to allow for the presence of structural changes in growth models is to use dummy variable regresors. Lastly, it would be better if future research provide some evaluation of our suggested policies, but this is outside the scope of this page.

Data Appendix

Y = Real GDP; *L* = Employment (Total economy); *K* = Net Capital Stock at 2000 prices (Total economy); *HKI* = Human Capital Index measured as average year of total schooling.

All data, excluding *HKI*, are taken and constructed from AMECO-EUROSTAT database. *HKI* (average year of total schooling, average year of primary schooling, average year of secondary schooling and average year of tertiary schooling) is retrieved from Barro and Lee (2010).

Total school enrolment rate is constructed from World Development Indicators (2011).

DUM74 dummy captures the impact of recession. It is computed as 1 from 1974-77, 0 otherwise.

DUM90 dummy captures the impact of recession. It is computed as 1 from 1990-91, 0 otherwise.

Shift96 dummy captures the impact of monetary policy shift (inflation targeting regime). It is computed as 1 from 1996-2008, 0 otherwise.

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