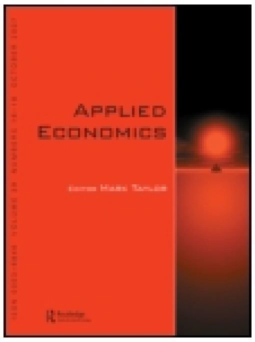

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The growth effects of education in Australia

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In this article, 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.

Keywords: steady state growth rate; economic growth; education; Australia

JEL Classification: C22; O56; O40

I. 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 (e.g. human capital) yield permanent growth effects. It started from 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 the Total Factor Productivity (TFP). However, an alternative approach is to extend 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 that can be classified depending on the type of human capital variables used by them. The first group links the output growth to some initial level or stock of educational attainment, such as school enrolment rates (e.g. Barro, 1991; Levine and Renelt, 1992; Mankiw *et al.*, 1992; Benhabib and Spiegel, 1994; Englander and Gurney, 1994; Loayza, 1994; Caselli *et al.*, 1996; Hauk and Wacziarg, 2004, among others). The second group relates growth to the flow of educational attainment rather than its level (e.g. 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 alternative measures of human capital based on both stocks and annual average growth rates (e.g. Gemmell, 1996; De La Fuente and Doménech, 2000, among others); they found the latter measure yields plausible estimates.

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

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

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attempts that used time series data include Jenkins (1995), Asteriou and Agiomirgianakis (2001), Loening *et al.* (2010) and Rao and Vadlamannati (2010). In the case of the UK, Jenkins (1995) found that highly qualified workers contribute almost twice to productive efficiency than 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 Gross Domestic Product (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, Loening *et al.* (2010) found that human capital (measured as the 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 article, 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 country is of particular interest because previous studies on earnings such as Miller *et al.* (1995) and Preston (1997) found high returns to advanced educational qualifications, and it is interesting to study empirically the impact of schooling on growth using different measures of education. To the best of our knowledge, no studies on long-run economic growth have examined the influence of schooling beyond the aggregate level. In addition, this is interesting because there are only a few studies that have estimated and analysed the SSGR using country-specific time series data. 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.[2] According to Bergheim (2008), enrolment rate is not a useful measure of human capital because it does not include information on years of education.[3] We show that alternative measures (total school enrolment rate, average year of 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 article is as follows. Section II discusses our extensions to the Solow model and develops our specifications. Empirical results are discussed and presented in Section III. Finally, Section IV concludes this article.

II. Specification

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

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

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

$$\Delta \ln y^* = SSGR = \Delta \ln A = TFP \tag{2}$$

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

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

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 nonlinear function of *HKI* (human capital measured as the average year of total schooling) as follows:

$$g = (\varpi HKI)t \tag{4}$$

$$g = (\varpi HKI)t + \gamma_1 HKI + \gamma_2 HKI^2 \tag{5}$$

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. Substituting Equation 5 into 3 in its intensive form gives

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

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

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

[2] 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.

[3] For example, let us 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% of 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% of secondary enrolment rate may not be sufficient to maintain the initial level of human capital. To this end, we need information about the 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.

Taking the first difference gives

$$\Delta a_t = TFP = \varpi_1 \Delta HKI_t \times T + \varpi_2 HKI_{t-1} + \gamma_1 \Delta HKI_t + \gamma_2 \Delta HKI_t \times HKI_t \quad (8)$$

Equation 8 can be interpreted as the intermediate period effects of HKI on $SSGR$.[4] In the long run, however, all the differences of the variables become zero in the steady state. Therefore, the $SSGR$ is

$$SSGR = \varpi HKI \quad (9)$$

Based on Equation 9, it could be asserted that the higher the HKI is, the higher the $SSGR$ becomes.

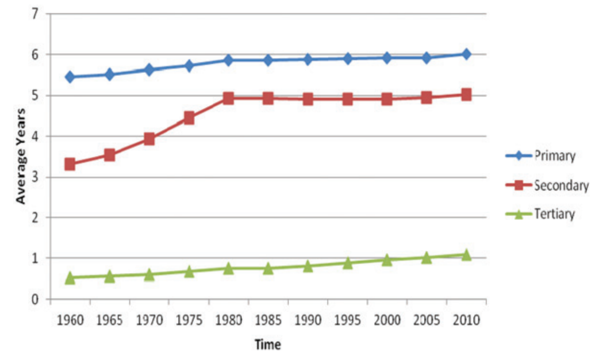

Fig. 1. Average years of primary, secondary and tertiary education in Australia

III. 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.[5] During 1960–2008, Australia has experienced an average GDP growth rate of 3.5%. Although during this period, Australia had encountered significant structural changes (e.g. 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[6]), its growth rate has remained well above 2.5% per year. The average growth rates over the sub-periods 1960–1969, 1970–1979, 1980–1989, 1990–1999 and 2000–2008 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. According to the Barro and Lee (2010) dataset, average educational attainment in Australia passed from 9.27 in 1960 to 12.12 in 2010. This huge increase in education probably helped to foster the GDP growth; this is the aspect we investigate in this study.

Figure 1 illustrates the average attainment with respect to primary, secondary and tertiary education.[7] 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 and 5 years. The average attainment in tertiary education has been the lowest and since 2000 it has reached 1 year. 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 1970s, retention rates in the secondary education dramatically increased followed by a sharp increase in enrolments in vocational colleges and universities. This is probably due to a rationalization of upper secondary school curriculum into a generic credential system which catered for all students, regardless of whether they intended to apply for a university place or not (Pascoe *et al.*, 1997). 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).

Figure 2 illustrates the evolution of HKI[8] 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 that have 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, it is interesting to study the role of education in explaining the $SSGR$ of Australia.

Unit root tests

Lee and Strazicich's (2003) two-break minimum Lagrange Multiplier (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.[9]

[4] 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.
[5] The Appendix contains data description and descriptive statistics (Table A1).
[6] 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.
[7] Barro and Lee's (2010) data is used to construct Figs 1 and 2.
[8] The dataset of Barro and Lee (2010) has observations for every 5 years between 1950 and 2010. Intermediate data are linearly interpolated (see, e.g. Bergheim, 2008; Park, 2010; 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.
[9] Model A: $Z_t = [1, t, D_{1t}, D_{2t}]'$ where $D_{jt} = 1$ for $t \geq T_{Bj} + 1, j = 1, 2,$ and 0 otherwise. Model C: $Z_t = [1, t, D_{1t}, D_{2t}, DT_{1t}, DT_{2t}]'$ where $DT_{jt} = t - T_{Bj}$ for $t \geq T_{Bj} + 1,\ j = 1, 2,$ and 0 otherwise. T_{Bj} denotes the break date. This technique involves deriving the LM test statistic (Kumar and Webber, 2013) and selecting the optimal lag lengths (Ng and Perron, 1995).

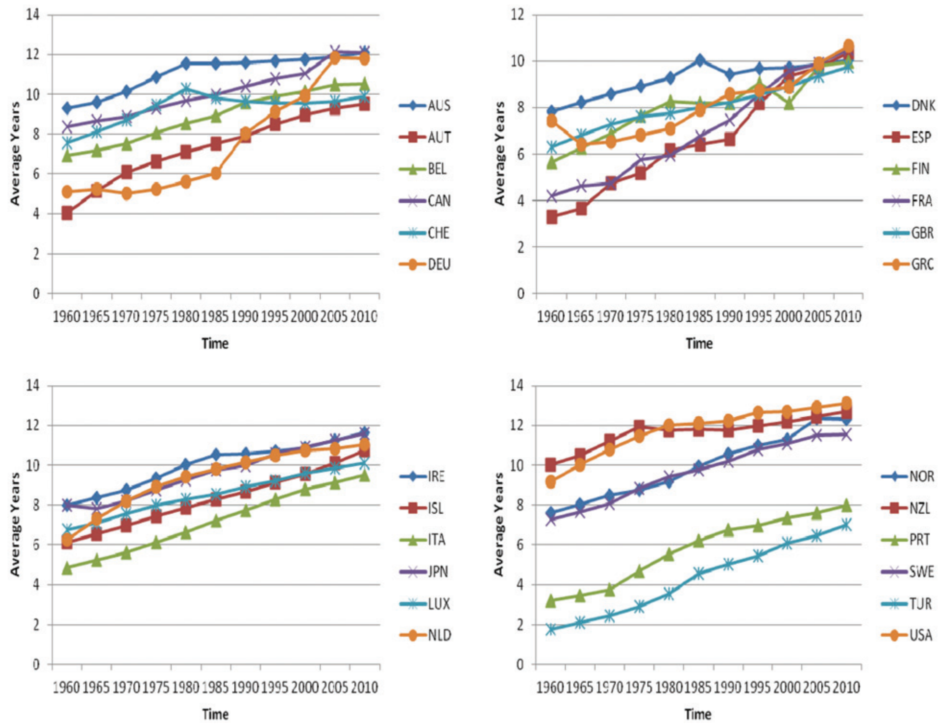

Fig. 2. Evolution of the average years of total schooling in advanced countries
Notes: AUS = Australia; AUT = Austria; BEL = Belgium; CAN = Canada; CHE = Switzerland; DEU = Germany; DNK = Denmark; ESP = Spain; FIN = Finland; FRA = France; GBR = Great Britain; GRC = Greece; IRE = Ireland; ISL = Iceland; ITA = Italy; JPN = Japan; LUX = Luxembourg; NLD = The Netherlands; NOR = Norway; NZL = New Zealand; PRT = Portugal; SWE = Sweden; TUR = Turkey and USA = United States of America.

Table 1. Two-break minimum LM unit root test, 1960–2008

	Level				First difference			
	Model A		Model C		Model A		Model C	
Variables	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

Notes: The 5% critical values for models A and C are −3.842 and −5.286, respectively. The numbers in square brackets indicate 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 (2003, 2004). Kumar and Webber (2013) contain more details on this test. RATS 7.2 was used to perform this test.

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

Estimates without structural changes

We utilized four techniques, namely Canonical Cointegrating Regression (CCR), General to Specific (GETS), Dynamic Ordinary Least Squares (DOLS) and Fully Modified Ordinary Least Squares (FMOLS) to estimate the extended

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 [5.125]***	−10.481 [5.808]***	−9.117 [2.779]**	−3.898 [2.337]**
α	0.323 [0.947]	0.298 [0.805]	0.673 [1.834]*	0.391 [0.864]
γ_1	1.580 [3.122]***	1.528 [3.259]***	1.465 [2.097]**	1.461 [2.354]**
γ_2	−0.075 [3.056]***	−0.073 [3.175]***	−0.068 [1.919]*	−0.070 [2.332]**
ϖ	0.001 [2.991]***	0.001 [2.828]**	0.001 [1.849]*	0.001 [2.066]**
λ	−0.368 [4.529]***			−0.397 [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

Notes: The *t*-statistics are in [] brackets. 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 heteroscedasticiy 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 Akaike Information Criterion (AIC). The SEs (not reported) for the DOLS estimation are calculated using the Newey–West correction. The GETS equation was estimated using nonlinear least squares as follows: (*r* squared was 0.41 and due to short sample only one lag was used)

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

All tests were performed using Eviews 7.0 software.
*, ** and *** denote significance at the 10, 5 and 1% levels, respectively.

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

$$\ln y_t = Intercept + \alpha \ln k_t + \gamma_1 HKI_t + \gamma_2 HKI_t^2 + \varpi HKI \cdot T \tag{10}$$

In the first instance, we estimated Equation 10 without allowing for any structural changes that were experienced in the domestic economy. Table 2 presents these results. Fairly consistent estimates were attained across the four estimators. The speed of adjustment (λ) implies negative feedback mechanism and is statistically significant at the 1% level. The Engle and Granger (1987) *t*-test supports the existence of cointegration among the variables at the 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 and 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 the 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) endogenous structural break tests. Since this test performs only when the parameters are linear, we utilize the OLS estimates of GETS for this purpose.[11] 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 the 1% level. The detected break dates are 1974 and 1996 and these are not unrealistic because Australia experienced a recession during 1974 and 1996, which 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

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

Table 3. Quandt–Andrews structural break tests, 1960–2008

Statistic	Value	Break date	Probability
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***

Notes: Probabilities calculated using Hansen's (1997) method, see Andrews and Ploberger (1994) for more details. Eviews 7.0 was used to perform this test.
***Indicates significance at the 1% level.

that time. Table 4 presents 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 a 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, the Wald statistic rejected the null of no break at the 5% level. Further, *F*-statistic rejected the null of no break at the 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's 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 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 ranging between 0.32 and 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 the 5% level. In Fig. 3, we present the actual and fitted values of $\Delta \ln y$ and the fit is satisfactory implying that the estimates are robust.

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 Fig. 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% found by us using the 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.

Estimates with alternative measures

Besides 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,[15] 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.[16] The GETS technique is used to attain these results. The magnitude of the growth effect of total school enrolment rate is 0.00038 and the average value of *SSGR* is around 0.3%, which seems very trivial.[17] Alternatively, the estimate of the growth effect of the average year of primary (tertiary) schooling is 0.00057 (0.00042) and to this end the *SSGR* is around 0.5% (0.4%). Moreover, the estimate of the growth effect of the average year of secondary schooling is 0.00030. With regard to the *SSGR* due to the average year of secondary schooling, it is very low at around 0.1%. In all cases, the estimates of the growth effects of *HKI* based on the three measures are statistically significant at the 5% level. From a comparative perspective, we argue that the average year of total schooling is the optimal measure of human capital and yields relatively higher value of *SSGR*.

IV. Conclusion and Policy Implications

This article 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 nonstationary

[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.
[14] Compared to the estimates of *SSGR* for the developing countries, estimates of *SSGR* for the advanced countries seem to be limited.
[15] Data on secondary and tertiary enrolment rates are not available. Therefore, total school enrolment rate is proxied by primary school enrolment (% gross).
[16] These results can be obtained from the authors upon request. The data on total school enrolment rate is retrieved from the World Development Indicators.
[17] Two dummies (1996 inflation-targeting regime and 1990 recession) were incorporated into the model and both were statistically significant at the 5% level.

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)

Notes: 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.

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 [13.291]***	−11.056 [13.303]***	−10.937 [6.203]***	−9.139 [5.282]***
α	0.315 [2.384]**	0.339 [2.031]**	0.429 [3.035]***	0.475 [2.154]**
γ_1	1.663 [8.165]***	1.624 [7.502]***	1.715 [5.362]***	1.801 [5.909]***
γ_2	−0.078 [7.860]***	−0.076 [7.218]***	−0.101 [4.930]***	−0.085 [5.751]***
ϖ	0.001 [5.260]***	0.001 [3.735]***	0.001 [3.495]***	0.001 [2.318]**
Shift96	0.054 [5.881]***	0.059 [5.125]***	0.057 [4.374]***	0.045 [3.429]***
DUM74	−0.025 [3.278]***	−0.023 [2.622]**	−0.029 [3.116]***	−0.027 [2.896]**
DUM90	−0.034 [3.518]***	−0.032 [2.660]**	−0.034 [4.997]***	−0.028 [2.217]**
λ	−0.766 [4.128]***			−0.789 [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

Notes: The *t*-statistics are in [] brackets. 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 heteroscedasticiy 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 SEs (not reported) for the DOLS estimation are calculated using the Newey–West correction. The GETS equation was estimated using nonlinear least squares as follows: (*r* squared was 0.46 and due to short sample only one lag was used)

$$\Delta \ln y_t = Intercept + \sum_{i=1}^{n1} \mu_{1i} \Delta \ln y_{t-i} + \sum_{i=1}^{n2} \mu_{2i} \Delta \ln k_{t-i} + \sum_{i=1}^{n3} \mu_{3i} \Delta HKI_{t-i} + \varphi_1 Shift96 + \varphi_2 DUM74 + \varphi_3 DUM90 + \lambda\left[\ln y_{t-1} - (Intercept + \alpha \ln k_t + \gamma_1 HKI_t + \gamma_2 HKI_t^2 + \varpi HKI \cdot T)\right]$$

All tests were performed using Eviews 7.0 software.
** and *** denote significance at the 5 and 1% levels, respectively.

and provided break dates that are located mostly during the 1980s and 1990s. Four time series techniques (CCR, GETS, FMOLS and DOLS) were utilized 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.[18] Second, we employed the Quandt–Andrews and the 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

[18] Except the DOLS estimate that is significant at the 10% level.

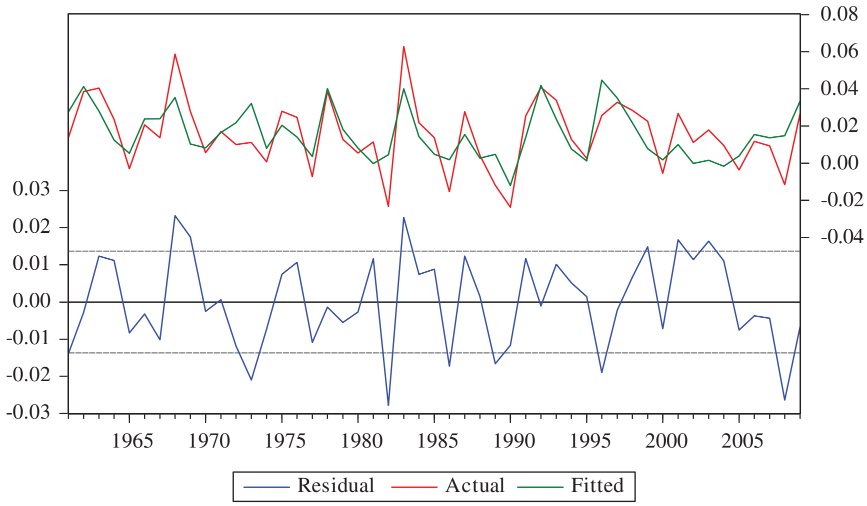

Fig. 3. Actual and fitted series of $\Delta \ln y$

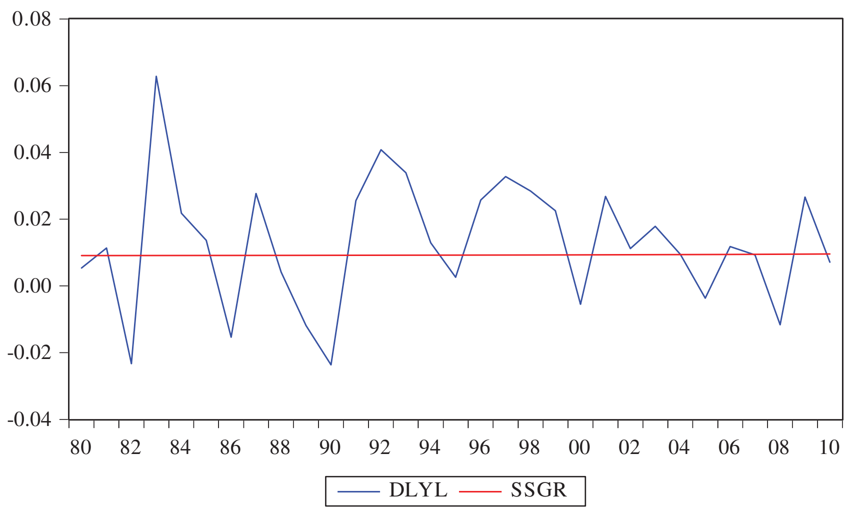

Fig. 4. *SSGR* for Australia
Note: $SSGR = 0.001^{*}HKI$

policy shift). Since the Chow method tests for exogenous breakpoints, we tested for a number of expected exogenous 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 the 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, namely 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 employed. The capital share is from 0.32 to 0.48 and has become statistically significant. More importantly, 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). Robustness test indicates that the average year of total schooling is the optimal measure of human capital and yields relatively higher value of *SSGR* for Australia.

From a policy perspective, the central question of interest is how educational attainment can be increased in Australia? It is well known that reforms are vital to improve educational attainment rate. In the case of 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. 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 Information and Communications Technology (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.

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Appendix: Data Description

Y = Real GDP; L = Employment (Total economy); K = Net Capital Stock at 2000 prices (Total economy); HKI = Human Capital Index measured as the 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 years of secondary schooling and average year of tertiary schooling) is retrieved from Barro and Lee (2010).

Total school enrolment rate (proxied by primary school enrolment (% gross)) is constructed from World Development Indicators (2011).

DUM74 dummy captures the impact of recession. It is computed as 1 from 1974 to 1977, 0 otherwise.

DUM90 dummy captures the impact of recession. It is computed as 1 from 1990 to 1991, 0 otherwise.

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

Table A1. Descriptive statistics

Variable	Mean	SD	Min	Max
y	0.063	0.013	0.040	0.085
k	0.171	0.038	0.121	0.251
HKI–average years of total schooling	11.095	0.974	9.296	12.119
HKI–average years of primary schooling	5.788	0.178	5.463	6.003
HKI–average years of secondary schooling	4.523	0.629	3.314	5.014
HKI–average years of tertiary schooling	0.783	0.192	0.518	1.101
HKI–total school enrolment rate	105.630	3.743	100.258	112.091

Notes: Min = minimum value and Max = maximum value. Data period for *HKI* – total school enrolment rate is from 1971 to 2008. For other variables, the data period is from 1960 to 2008.