Testing the balanced growth hypothesis: Evidence from China

Hong Li* and Vince Daly, Kingston University

Abstract

We investigate whether China's experience during 1952-2004 supports the balanced growth entailment of the neoclassical growth model. Estimation of long-run relations among output, consumption and investment for the full period reject the balanced growth hypothesis for both the national and regional economies. When the economic reforms of the late 1970s are modelled as a structural break by the methods of Johansen *et al*. (2000) and Perron (1989), we find some evidence of balanced growth in the pre-break period but in the post-break period the 'great ratios' are trend-stationary, precluding fully balanced growth, though permitting a common (stochastic) productivity trend.

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* corresponding author. We are grateful for advice received from colleagues.

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1 Introduction

Kuznets' (1942) study of the macroeconomic aggregates of the USA during that country's period of industrialisation led him to posit a long-run constancy in the ratio of savings to income. Klein and Kosobud (1961) applied more formal trend-fitting methods to Kuznets' data and concluded that some of the 'great ratios' were constant but others, including savings/income, actually possessed a slight trend. At the same time Kaldor (1961) posited a number of constancies, though not including the savings/income ratio, as "stylised facts" of the growth process. These empirical observations helped to launch what has been called the "balanced growth" literature, characterised in King *et al.* (1991, p819) as claiming empirical support for '… balanced growth in which output, investment and consumption all display positive trend growth but the consumption:output and investment:output "great ratios" do not ....'

This paper investigates the extent to which the balanced growth hypothesis is applicable to the development of the Chinese economy since the middle of the last century. During this period the political institutions have been such that the central government of China has been able to initiate radical structural change; we focus particularly upon the reduction of central planning that was initiated in the late 1970s. Following these economic reforms, China has sustained high economic growth – in marked contrast to the pre-reform decades. Overall, this is a qualitatively different context to that of the USA, which provided an empirical underpinning to the initial balanced growth propositions. We examine whether such propositions remain applicable if transferred from the USA experience to that of China. The question is of more than purely academic interest since the presence of balanced growth speaks for, rather than against, the relevance of the neoclassical growth model – which is a model offering relatively little opportunity for government to play a significant role in the growth process.

The neoclassical growth model predicts a balanced growth path along which per capita output, consumption and investment grow at the same constant rate while the investment/output and consumption/output ratios are constant. This hypothesis has not been consistently supported by empirical investigation using data from developed countries. King *et al*. (1991), using cointegration analysis, find that post-war U.S. macroeconomic data exhibit a common stochastic trend, satisfying parameter restrictions consistent with the balanced-growth hypothesis. Mills (2001) uses the technique of generalised impulse response functions as well as Johansen's method for estimating cointegration rank, and finds evidence to support the stationarity of the 'great ratios' for the UK during the post-war period. On the other hand, Kunst and Neusser (1990), using Johansen's method, strongly reject the hypothesis of stationary 'great ratios' for Austrian data. Serletis (1994) does not find any evidence of stationary ratios in multivariate analysis of Canadian data (1929-1983). Furthermore, Serletis and Krichel (1995) and Harvey *et al*. (2003) do not find evidence from OECD and G7 countries, respectively, to support the hypothesis of balanced growth. Importantly, from the perspective of this paper, Clemente *et al*. (1999) argue that the evidence against balanced growth is less convincing when the possibility of structural breaks is taken into account.

In the light of the mixed empirical results in the literature, we are motivated to examine the empirical support for the balanced growth hypothesis in a developing country, in particular

China, as opposed to the advanced industrialised countries. The case of China is of particular interest due to the noticeable change in the pace of economic growth following economic reforms and also because of the diversity in the regional economies. The regime switch and the lack of uniform development across China's regions (Li and Daly, 2005), both allow us to explore whether the evidence for balanced growth varies with the level of development.

This empirical study is underpinned by the neoclassical stochastic growth theory developed by Brock and Mirman (1972), Donaldson and Mehra (1983) and implemented empirically by King *et al*. (1988). Following King *et al*. (1991), Serletis (1994) and Serletis and Krichel (1995), we investigate the Chinese growth process via multivariate analysis of major macroeconomic aggregates and also univariate analysis of selected 'great ratios'. Specifically, we apply Johansen's (1995) maximum likelihood approach to estimate the number of long-run steady-state relations among Chinese per capita output, consumption and investment, at the national and regional levels and to test the validity of the parametric restrictions that characterise balanced growth. Our novel contribution to the empirical testing of balanced growth is that we extend the cointegration analysis to allow for structural change by using the method of Johansen *et al*. (2000). We further use Perron's (1989) approach to incorporating a structural break in univariate unit-root testing, allowing us to assess whether the log ratios of consumption vs. output and investment vs. output are stationary – as is required by the hypothesis of balanced growth. The following questions will be examined. Are the time series properties of Chinese national and regional per capita output, consumption and investment properties consistent, in whole or in part, with the balanced-growth predictions of the neoclassical growth model? Does the level of development affect the evidence for balanced growth? Are the conclusions sensitive to the inclusion/exclusion of a structural break to recognise the major economic reforms of the late 1970s?

The paper is organised as follows. Section 2 summarises the neoclassical growth theory and its econometric representation; section 3 describes the data and conducts a preliminary analysis of their time series properties; section 4 applies the standard Johansen framework to determine whether the Chinese macroeconomic variables possess cointegrating vectors in number appropriate for balanced growth and whether these vectors additionally satisfy the parameter restrictions required for stationarity of the 'great ratios'. Section 5 extends the analysis by using the methods of Johansen *et al.* (2000) and Perron (1989) to take into account the possibility of a structural break caused by the economic reforms of the late 1970s. Section 6 concludes.

2 Theoretical considerations and econometric representation

The Solow and Swan (non-stochastic) growth theory for a one-sector economy specifies a neo-classical production function exhibiting constant returns to scale and diminishing marginal product with respect to each input. The model assumes that technology augments the labour input to an extent which grows at an exogenously fixed rate, g. Fixed capital formation is provided by a constant fraction, s, of unconsumed output. Production technology and consumer preferences are assumed to be such that a steady state solution exists. The steady-state per capita quantities: per capita capital, k; per capita output, y; and per capita consumption, c, then all grow at the same rate as technological progress. Their aggregate levels, K, Y and C, grow accordingly in the steady state at a common rate of $g+n$, where n is an exogenously determined rate of growth for the size of the labour force.

In the particular case of the Cobb-Douglas production function we have

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

A_t represents exogenous labour-augmenting technical progress. The constant growth rate for the effectiveness of labour can be represented as a deterministic logarithmic trend: $\log(A_t) = g + \log(A_{t-1})$. The per capita variables all grow at the exogenously given rate of technical progress: $\gamma_y = \gamma_k = \gamma_A = g$ ($\gamma_x \equiv \log(x_t) - \log(x_{t-1})$, $x_t = X_t / L_t$). This "balanced growth" of the per capita aggregates implies that the 'great ratios' of investment and consumption to output are constant.

Alternatively, technological progress can be modelled stochastically (King *et al.*, 1988) as a logarithmic random walk with drift:

$$\log(A_t) = g + \log(A_{t-1}) + \varepsilon_t \tag{2.}$$

The term, ε_t, is a white noise process representing productivity shocks, i.e. deviations of the growth rate for labour effectiveness from its expected value, g. Each ε_t contributes a permanent impact: their cumulative effect at any point in time is their undiscounted sum up to and including that date, which is a 'stochastic trend'. In the solution path for a log-linear version of the stochastic growth model the logarithms of output, consumption and investment consequently all share this stochastic trend (King *et al.*, 1988). These three aggregates are therefore individually non-stationary, being integrated of first order - $I(1)$. Nevertheless, because they have in common a single stochastic trend, it is possible to construct two distinct 'cointegrated' linear combinations of them that are stationary. The model solution further implies that these combinations are the logarithms of the consumption / output and investment / output 'great ratios', i.e. there is balanced growth in a stochastic sense. The testable implications of the model then include (1) an appropriate cointegration rank and (2) parametric restrictions to ensure that the cointegrating vectors imply stationarity of the 'great ratios'. The required cointegration rank, r, for an error-correction model in the logarithms of per capita output, consumption and investment is $r = 2$. For the cointegrating vectors to imply "balanced growth", their normalised coefficients should be $(1 \quad -1 \quad 0)$ and $(1 \quad 0 \quad -1)$, and there should be no trend in the cointegration space. These restrictions can be tested within the Johansen (1995) framework or alternatively by directly assessing the stationarity of the 'great ratios'.

3 Data: sources and characteristics

The raw data used in this study are those of GDP, final consumption expenditure, gross fixed capital formation, consumer price index and population between 1952 and 2004, provided at provincial level in The Gross Domestic Product of China 1952-1995[1] and various issues of the Statistical Yearbook of China published by the State Statistical Bureau of China. We combine Sichuan and Chongqing[2] as a single province and omit Tibet due to its incomplete data availability. This yields data series for 28 provinces, at annual frequency, which are then aggregated to regional level, as described below, to compile the variables of interest.

The variables of interest in this study are real per capita GDP, real per capita consumption expenditure and real per capita fixed capital formation at national and regional levels. These variables are constructed as follows. First, the provincial series for GDP, final consumption

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[1] This is compiled by the Department of National Economic Accounting, State Statistical Bureau and published by Dongbei University of Finance and Economics Press

[2] Chongqing was separated as a province from Sichuan in the early 1990s.

expenditure and gross fixed capital formation are converted from nominal to real units by division by their respective provincial consumer price index. Regional series for the major administrative regions - Eastern, Central and Western, are then formed by aggregation across the provinces assigned to each region. National series are produced by aggregation of the provincial series directly. Similar regional and national aggregation is performed upon the provincial population data in order to construct series for real per capita output (y), real per capita consumption (c) and real per capita investment (i) at national and regional levels.

In this section, we examine the time series properties of these per capita series at the national and regional levels. Figure 1 shows that over time the three variables, in logarithms: *logy*, *logc* and *logi*, have increased at both national and regional levels of aggregation. An upward trend had become obvious since the late 1970s, prior to which the series were more volatile and possibly tended to decline. On the basis of graphical inspection, the three aggregates seem to share similar trend tendencies. We would like to test more formally whether or not the series share a common stochastic trend, as is implied by the balanced growth conclusions of the stochastic neoclassical growth model.

[Figure 1 is about here.]

To implement the formal cointegration tests within the Johansen framework, we first need to establish that the national and regional *logy*, *logc* and *logi* are $I(1)$ series. We use the augmented Dickey-Fuller test, including a constant and trend in the test equations. The lag length is determined using downward testing beginning with an arbitrarily large number of lags, in this case 10, on the basis of a modified Akaike Information Criterion (Ng and Perron, 2001).

[Table 1 is about here.]

The test results presented in Table 1 indicate that these series may be treated as $I(1)$, with the exception of *logc* in the Eastern region and nationally – which show evidence of being $I(2)$. We proceed nevertheless, because $I(2)$ behaviour of consumption does not rule out partially balanced growth, with stability of the investment/output ratio, and also because of our interest in the possibility of structural breaks. (This apparent $I(2)$ behaviour may be seen as signalling a structural break in the Eastern region – where post-liberalisation growth has been relatively rapid, and consequently nationally.) Having established that the per capita variables show evidence of possessing stochastic trends, we move to consider below whether these are idiosyncratic or shared in common – the latter being a necessary condition for "balanced growth".

4. Testing for balanced growth without a structural break

In the absence of cointegration, *logy*, *logc* and *logi* – for a particular region or nationally, are driven by three separate stochastic trends, precluding balanced growth. We use Johansen's (1995) approach to estimate the number of cointegrating vectors constraining the long-run behaviour of these series. The existence of two cointegrating vectors involving three series would imply that the series share in common a single stochastic trend, introducing the possibility of balanced growth. The results of this cointegration analysis are reported in Table 2. Because the presence of a linear trend within a cointegrating relationship would rule out balanced growth, we use the test variant in which linear time trends are permitted within the data series but not within the cointegration space. Lag lengths in the test VECMs were selected by sequential modified LR test statistics and found to be 2 or 3 for all cases.

[Table 2 is about here.]

Cointegration analysis suggests that *logy*, *logc* and *logi* share a common stochastic trend in the Central region and also in the Western region, satisfying a necessary, but not sufficient, condition for balanced growth in those regions. Nationally, however, and in the Eastern region the standard Johansen procedure suggests a single cointegrating vector, which imposes insufficient constraint upon the long run behaviour of the three time series for *log(c/y)* and *log(i/y)*, which are the 'great ratios' in logarithmic form, to both be stationary. This result should be interpreted in the context of the previous diagnosis that *logc* ~ *I(2)* at the national level and in the Eastern region. As is argued more completely in Everaert (2003), an *I(2)* series – here *logc*, cannot form a stationary combination with two *I(1)* series – here *logy* and *logi*. The cointegration space is then two-dimensional, rather than three-dimensional, so that the cointegration rank must be zero or unity to be consistent with the previous diagnosis that *logy*, *logi* are *I(1)* series. A cointegration rank of unity for Eastern and national data then leaves open the possibility that one 'great ratio', *log(i/y)*, is stationary but not both.

It is not surprising to discover similar diagnoses of series characteristics at the national level and in the Eastern region since the Eastern region accounts for the largest proportion of the Chinese economy. The diagnosis of *I(2)* behaviour in the Eastern and national *logc* might be evidence of a permanent feature, such as a drift in the average propensity to consume, that precludes balanced growth. Alternatively, it might point to the existence of a structural break in the series, with *logc* ~ *I(1)* on either sided of the break and leaving open the possibility of balanced growth before and/or after the break. As is well known, the Eastern region was selected to lead the process of economic development following the market liberalisation reforms of the late 1970s. The Eastern region has consequently experienced the most radical structural break between the first and second halves of the data period and it may be that this structural break is mis-interpreted by the standard Johansen procedure as a second unit root process in the Eastern region, hence also at the national level. The possibility of modifying the standard procedure to explicitly incorporate a structural break within the cointegration analysis will be investigated in the next section.

Now we examine the extent of empirical support for the further implication of the neoclassical stochastic growth theory that the 'great ratios' should be stationary stochastic processes. This can be interpreted, within the Johansen framework, as a requirement that the normalised coefficients of the cointegrating vectors for *logy*, *logc* and *logi* should be $(1 \quad -1 \quad 0)$ and $(1 \quad 0 \quad -1)$. We carry out likelihood ratio tests to assess whether these restrictions are acceptable, individually or simultaneously, given the cointegration ranks reported in table 2. The results are presented in Table 3.

[Table 3 is about here.]

Table 3 shows that, at national level and in all three regions, the data does not support the parameter restrictions required for stationarity of *log(c/y)*. In this sense, the balanced growth hypothesis is rejected even where the cointegration rank is appropriate to it.

Balanced growth, in the sense of the stationarity of *log(c/y)* and *log(i/y)* could also be assessed outside of the Johansen framework by directly investigating the stationarity of these two series via unit-root testing. ADF test results (available on request) agree with the parameter restriction tests in that they suggest non-stationarity for both of these series in all three regions. At the national level the unit root null hypothesis is rejected for both series if the ADF test equation is permitted to include a linear trend, but such a trend then implies deterministic non-stationarity for the series being tested and thus still precludes balanced growth in the strict sense of constant means for the 'great ratios'.

5. Testing for balanced growth in the presence of a structural break

In section 4, we found that Johansen's method rejects balanced growth, nationally and in all three regions individually - either because the macroeconomic series do not share a common stochastic trend, or, where they do so, because the cointegrating vectors do not satisfy the required parameter restrictions. Direct unit root testing of the 'great ratios' gave no grounds to question this overall dismissal of the possibility of balanced growth.

We noted in passing that the economic reforms of the late 1970s could be seen as implying a parametric structural break in a VECM representation of the data series. Such a structural break can undermine the reliability of statistical methods that do not admit its presence. Accordingly, in this section, we employ alternative statistical methods that allow us to include a structural break in the analysis, in case this might change the conclusions with respect to the existence of balanced growth in the pre-reform and/or post-reform periods.

The graphs (Figure 1) of the national and regional log per capita consumption, output and investment offer visual support for the possibility of a break in trend behaviour around 1978, when the economic reforms started. We use the method developed by Johansen *et al.* (2000)

to estimate the cointegration rank in the presence of broken linear trends in the individual series and additionally use Perron's (1989) method to test for the stationarity of *log(c/y)* and *log(i/y)* in the presence of such a structural break.

Johansen *et al.* (2000) divide a sample of T observation points into q sub-periods by exogenous selection of break-points, T_j ; $j = 0, 1, 2, \cdots q$, with $T_0 = 1$ and $T_q = T$. The maintained model is a levels VAR of order k, hence a VECM of order $k\text{-}1$, in a p-vector of jointly endogenous variables. The first k observations of each sub-period are reserved as initial values, so that for each instance of $j = 1,2,\cdots q$ the observation points $T_{j-1} + k + 1 \le t \le T_j$ are described by

$$\Delta X_t = \begin{bmatrix} \Pi & \pi_j \end{bmatrix} \begin{bmatrix} X_{t-1} \\ t \end{bmatrix} + \mu_j + \sum_{i=1}^{k-1} \Gamma_i \Delta X_{t-i} + \varepsilon_t \qquad (3.)$$

Only the p-vectors of parameters relating to the deterministic components - π_j , μ_j, may vary between sub-periods. Cointegration is represented, as in the standard Johansen framework, by the restriction $\Pi = \alpha\beta'$, where α and β are $p \times r$ matrices. The possibility of quadratic trends for the levels series is eliminated by the additional maintained assumption that $\pi_j = \alpha\gamma_j$ in each sub-period.

The models for the separate sub-periods can be written jointly by defining dummy variables:

$$D_{j,t} = \begin{cases} 1 \; iff \;\; t = T_{j-1} \\ 0 \; otherwise \end{cases} \quad , \quad j = 1,2,\cdots q$$

and

$$E_{j,t} = \sum_{i=k+1}^{T_j - T_{j-1}} D_{j,t-i} = \begin{cases} 1 \; iff \; T_{j-1} + k + 1 \le t \le T_j \\ 0 \; otherwise \end{cases}, \quad j = 1,2,\cdots q$$

Each $D_{j,t}$ is an indicator for the end of the *(j-1)*th sub-period and its lagged values therefore point to the individual observation points within the j^{th} sub-period. By summing these lagged values $E_{j,t}$ acts as an indicator dummy for an entire sub-period. Using these dummy variables, the separate sub-period models can be combined into one as follows.

$$\Delta X_t = \alpha \begin{bmatrix} \beta \\ \gamma \end{bmatrix}' \begin{bmatrix} X_{t-1} \\ tE_t \end{bmatrix} + \mu E_t + \sum_{i=1}^{k-1} \Gamma_i \Delta X_{t-i} + \sum_{i=1}^{k} \sum_{j=2}^{q} \kappa_{j,i} D_{j,t-i} + \varepsilon_t \qquad (4.)$$

Here μ contains the intercept coefficients for the q sub-periods, and the terms $\kappa_{j,i} D_{j,t-i}$ are included to ensure that the estimation of the parameters of interest is not influenced by data at the initial observation points within each sub-period.

The time plots in Figure 1 are visually supportive of the possibility that the Chinese series experienced a break in trend at 1978, associated with the significant reforms to economic institutions. Since 1979, GDP, consumption and investment appear to have been trending upwards, in contrast to the preceding years. There is some visual evidence of other disruptive events, most notably the 'Great Leap Forward' of 1958-61, but since our focus is on the possible impact of the economic reforms that commenced around 1978, we set q=2, with T_0=1952, T_1=1978 and T_2=2004.

The cointegration rank is estimated by a variant of the canonical correlations approach of Johansen (1995, Ch.6). Johansen *et al.* (2000) present a table, based on response surface analysis, which enables the investigator to select appropriate parameters for a gamma distribution to be used for setting the critical values of the relevant test statistics. The parameters of the gamma distribution vary according to the number and location of sample break-points; in our case, with a single break-point, $v = T_1/T_2 = 0.4717$ is a key ratio for determining these parameters. Our results are reported in Table 4.

[Table 4 is about here.]

According to the standard analysis summarised in Table 2, the cointegration rank of the national series for *logy*, *logc* and *logi* is too small to support balanced growth, whether the nominal size for the relevant testing is set at 5% or 10%. Table 4 shows that, with nominal test size at 5%, this remains the case when the deterministic components of the individual series are permitted a structural break at 1978. At the weaker 10% significance level, however, the national series do now show a cointegration rank appropriate to balanced growth. At the regional level, the standard analysis found a cointegration rank appropriate for balanced growth in the Western and Central regions. Now, with a structural break at 1978, this is true only of the Western region.

As was noted previously, balanced growth requires not only an appropriate cointegration rank but also parameter values in the cointegrating vectors that are consistent with stationarity of the 'great ratios'. Figures 2 and 3, plotting *log(c/y)* and *log(i/y)*, respectively, at the national and regional levels, do not offer clear visual evidence of such stationarity, before or after 1978.

[Figures 2 and 3 are about here.]

For a formal assessment of the stationarity of the great ratios in the presence of a structural break at 1978, we follow Perron (1989) which modifies the ADF test procedure to allow an exogenously dated break in the deterministic component of a series. Specifically, the test equation is

$$\Delta y_t = \mu + \beta t + \delta_1 DU_t + \delta_2 DT_t + \delta_3 D_t + \alpha y_{t-1} + \sum_{i=1}^{i=k} \gamma_i \Delta y_{t-i} + \varepsilon_t \qquad (5.)$$

The dummy variables D_t, DU_t and DT_t are defined by reference to the break date, $t=T_B=1978$, viz:

$DU_t = 1$ for $t > T_B$, otherwise $DU_t = 0$;

$DT_t = t$ for $t > T_B$, otherwise $DT_t = 0$;

$D_t = 1$ for $t = T_B + 1$, otherwise $D_t = 0$.

The null hypothesis is that the individual 'great ratio' has a stochastic trend generated by a unit root process which may experience a pulse and a change of the drift rate at the break. This null imposes $\alpha=0$ ("unit root"), $\beta=0=\delta_2$ (no deterministic trend[3] before or after the break), and permits $\delta_1\neq0$ (a change of drift rate), $\delta_3\neq0$ (a pulse). The alternative hypothesis is of stationary fluctuations about a possibly broken deterministic trend, i.e. $\alpha<0$ ("error correction"), $\delta_3=0$ (no pulse), and permitting $\delta_1\neq0$ ("change of intercept"), $\delta_2\neq0$ ("change of slope"). The test is conducted on the basis of the estimated value of the autoregressive parameter, α, using critical values in Table VI.B of Perron (1989). Balanced growth requires

[3] Because the unit root would integrate this into a quadratic drift.

rejection of the null hypothesis: $\alpha=0$, in favour of $\alpha<0$, together with evidence that the deterministic trend is insignificant. Under these conditions, the 'great ratios' are mean-reverting and their means are constant either side of the structural break. Their means may have been permanently shifted by the structural break at 1978 but shocks other than this break have only a temporary effect.

To determine the lag length in the test equation (5), we employ the 't-sig' method described in Perron (1994) and commonly used in empirical studies dealing with trend-break unit root tests. Specifically, we start with the maximum lag length, ten in this case, and reduce this, one step at a time, until the coefficient of γ_k is significant at 10%. Table 5 reports the results of the unit root test by Perron's method for the log ratios at the national and regional levels with a break in 1978.

[Table 5 is about here.]

Table 5 indicates that we can generally reject the unit root hypothesis ($\alpha=0$) for national and all regional 'great ratios' in favour of the alternative hypothesis of trend-stationarity throughout 1952-2004, provided that a structural break is permitted in 1978. This rejection is at significance levels of 5% or better in all cases except for Eastern *log(i/y)*, where it is not much weaker than 5%. For such trend-stationarity to imply balanced growth it is additionally necessary that the trend slope be zero. Wald tests (not reported) reject the joint hypothesis $\beta=\delta_2=0$, i.e. "zero trend in the great ratios both before and after 1978", except for Central *log(i/y)*. In Table 5 we report the results of testing for zero trend separately in the two sub-periods periods before and since 1978. The hypothesis of zero trend is expressed in the pre-break period as $\beta=0$ and after the break as $\beta+\delta_2=0$. Given the rejection of unit roots, we use standard test procedures to test these restrictions. The t-statistics associated with the estimation of β suggest that the great ratios are stationary prior to 1978, nationally and in the regions, with the exception of a negative trend for *log(c/y)* in the Eastern region and a positive trend for *log(i/y)* in the Western region. In contrast, for the post-break period, Wald tests of $\beta+\delta_2=0$ suggest that non-zero linear trends are present in both series nationally and for all regions following 1978. The rejection of zero post-break trend is emphatic, being marginal only for *log(i/y)* in the Central region.

Introducing a structural break to represent the initiation of the major economic reforms of the late 1970s has substantially altered our conclusions about the extent of evidence favouring the balanced growth hypothesis. Firstly, with regards to the existence of a shared common trend for the macroeconomic aggregates, persisting throughout the complete period of observation, we now find some support for this nationally, but less support than previously at the regional level. As to the further requirement of a constant mean for both of the great ratios, of which we previously found no instance at all, we now find evidence for such in the pre-break period at the national level and in the Central region, with the other regions showing such constancy for one ratio but not for both. The non-stationarity of the Eastern *log(c/y)* and the Western *log(i/y)* in the pre-reforms period may be a consequence of the Chinese government's development strategy, which aimed to address issues of inequality at the expense of immediate efficiency in the allocation of support for investment projects. The (richest) Eastern region received relatively little public investment, so that the decline of its consumption/income ratio may reflect a need to fund investment through private saving. At the same time, the (poorest) Western region received additional public investment funding via the so-called Third Front strategy of the mid-1960s, reflected in a positive trend for its investment/income ratio.

In the post-reform period, the 'great ratios' are universally trend-stationary, with the trend ($\beta+\delta_2$) being universally negative for *log(c/y)* and positive for *log(i/y)*. This might be seen as reflective of the growth-oriented strategy implemented together with the economic reforms of the late 1970s. Since 1978, the geographical advantage of the coastal region has been used to integrate China with the outside world. The Eastern region, hence the nation as a whole, has benefited from increasing inflows of foreign direct investment while the Western region has received increasing public investment following the draft of the Western Development Strategy in the late 1990s. It is arguably the case that the public investment and the foreign direct investment are both features of late twentieth century China to a much greater extent than was the case for the USA a century earlier, whose data underpinned the first suggestions of constancy in the great ratios.

6 Conclusion

Motivated by the stochastic variant of neoclassical growth theory, this study has looked for evidence of balanced growth in the Chinese per capita output, consumption and investment at the national and regional levels. We are interested to know whether the conclusions are sensitive to the inclusion / exclusion of a structural break in 1978 and whether there might be some association between the level of economic development and the existence of balanced growth.

We find that the extent of evidence to support the balanced growth hypothesis does indeed depend noticeably upon whether or not the statistical procedures explicitly recognise the economic reforms of the late 1970s as a structural break. Balanced growth requires both an appropriate cointegrating rank for the macroeconomic aggregates, to provide these aggregates with a common stochastic trend, and also appropriate restrictions upon the cointegrating vectors to ensure stationarity of the 'great ratios'. As to the existence of common trends, including a structural break leads to some (weak) evidence of a common trend in the national data, where this was otherwise lacking, and alters the conclusions regarding common trends in the regions separately. Conclusions regarding the stationarity of the 'great ratios' are sensitive to the choice of test procedure and the different procedures produce conflicting evidence on occasion. In the case of unit-root testing, which offers no support for balanced growth when the potential structural break is ignored, we find that the inclusion of a structural break leads to contrasting conclusions in the pre-break and post-break periods. In the pre-break period, both 'great ratios' are assessed as stationary at the national level and at least one of them in every region – fully supporting the balanced growth hypothesis prior to 1978, for the national economy and for the Central region. In the post-break period both 'great ratios' are diagnosed to have non-zero trends, precluding balanced growth in the strict sense, in all sets of data series.

In considering whether balanced growth is more likely at different stages of economic development, we note that the variations between regions in the extent of evidence for balanced growth does not associate in any simple way with regional variations in the extent of economic development. We do, however, find that stationary 'great ratios', which are an entailment of balanced growth, are present to some extent before the economic reforms, but completely absent in the context of the increased extent and pace of development that has followed those reforms.

In the post-break period both 'great ratios' are trend-stationary, i.e. mean-reverting but with non-constant means, nationally and in all regions. This may be seen as indirect evidence of a common stochastic trend for the macroeconomic aggregates since 1978, where the method of Johansen *et al*. (2000) found only weak evidence of such common trends being present throughout the entire span of the data. It could be argued therefore that although, we have found that the strict requirements of balanced growth are not empirically supported by China's recent growth experience, we have found some regularity amongst the macroeconomic aggregates which echoes Klein and Kosobud's (1961) discovery of steady trends in some of the great ratios during an earlier period of industrialisation in the USA.

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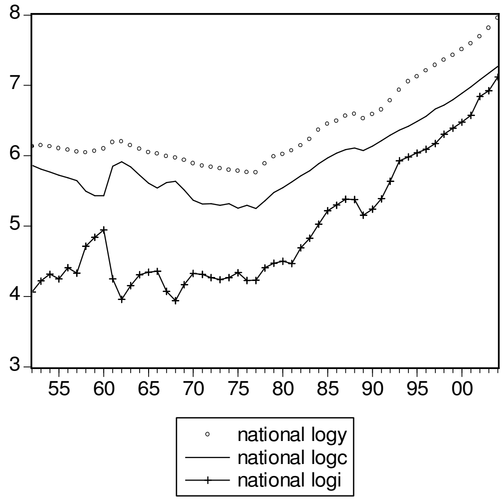
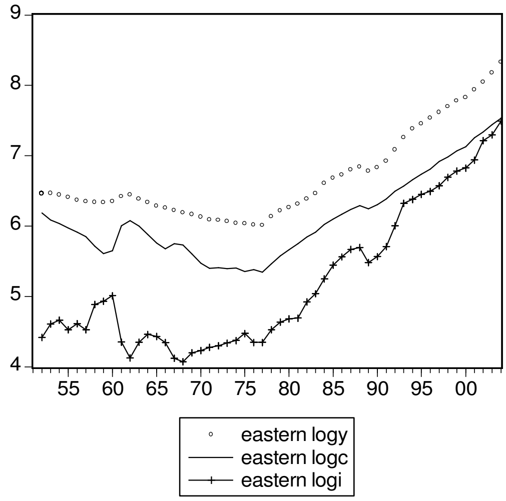
Figure 1 *logy*, *logc* and *logi* at the national and regional levels during 1952-2004

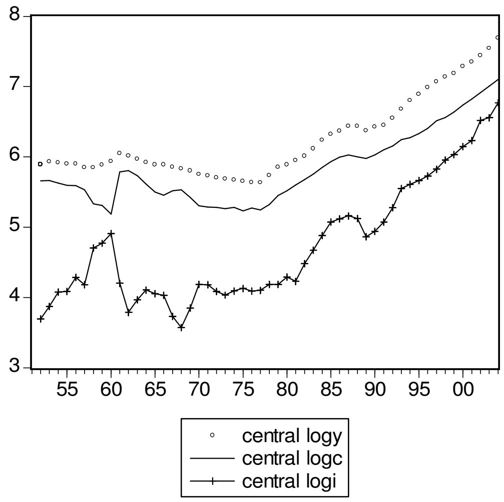
National



Eastern Region



Central Region



Western Region

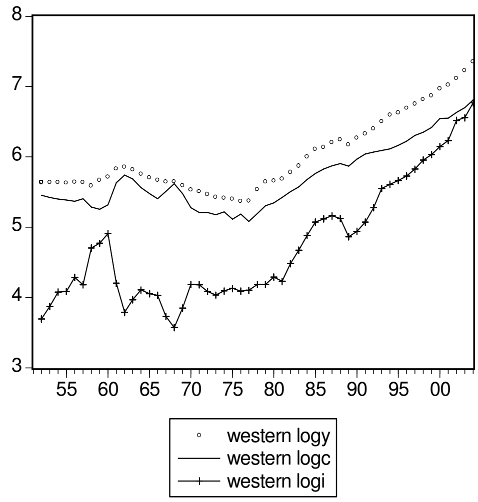


Figure 2 Log ratio of consumption to output at national and regional levels, 1952-2004

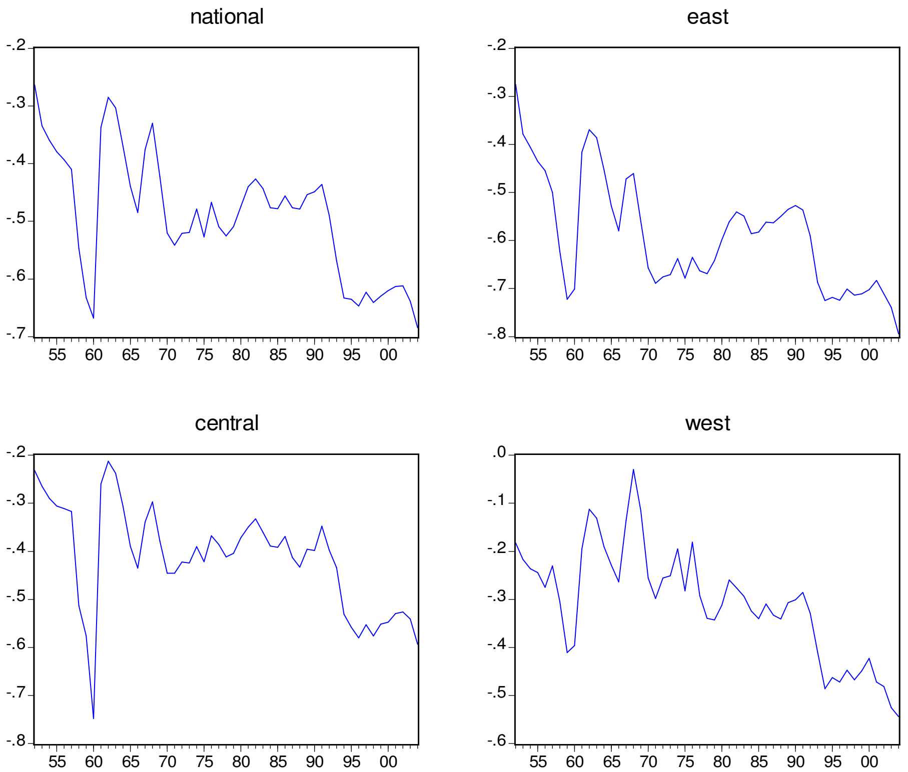


Figure 3 Log ratio of investment to output at national and regional levels, 1952-2004

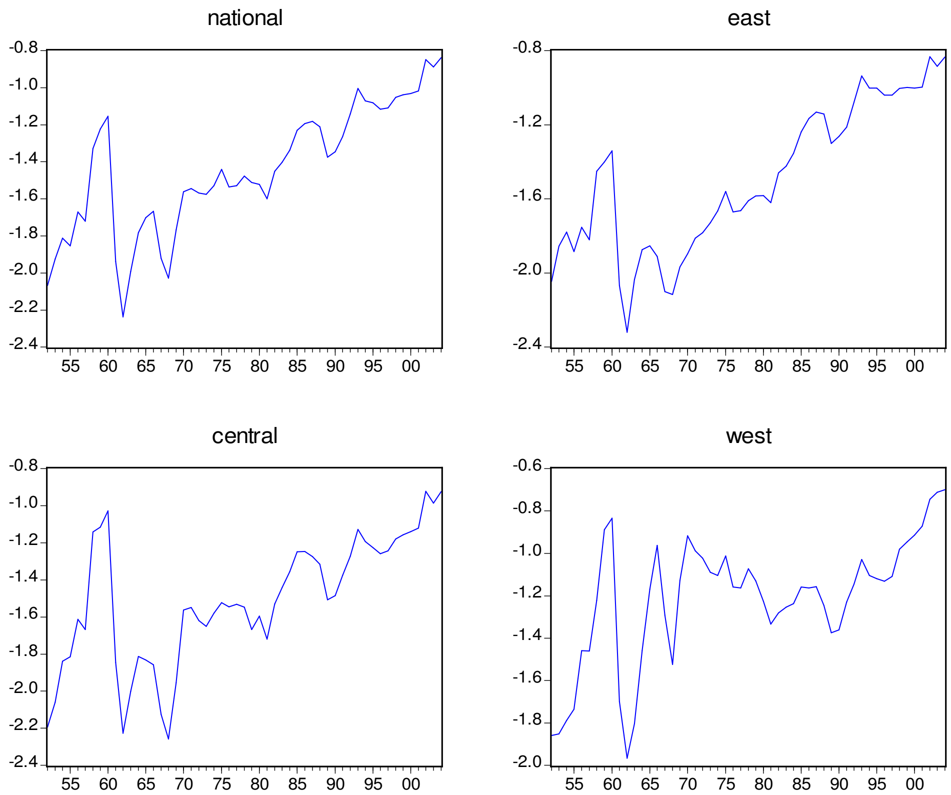


Table 1 ADF tests on *logy*, *logc* and *logi*, 1952-2004

Region	*Logy*		*Logc*		*Logi*	
	Lag	ADF	Lag	ADF	Lag	ADF
a) series levels						
National	2	-0.116	3	-0.682	2	-0.869
Eastern	2	-0.083	3	-0.886	0	-0.771
Central	1	-0.371	3	-0.825	2	-1.258
Western	1	-0.393	3	-0.533	4	0.1075
b) First differences						
National	0	-3.598**	5	-2.422	0	-5.786***
Eastern	0	-3.760**	5	-2.246	0	-6.136***
Central	0	-4.013**	0	-6.775***	0	-5.759***
Western	0	-4.341**	0	-5.210***	0	-5.458***

Note: *, ** and *** denote rejection of the unit root null with significance level at 10%, 5% and 1%, respectively.

Table 2 Johansen cointegration test on *logc*, *logy* and *logi*, 1952-2004

	Hypothesized No. of CE(s)	Trace Statistic	p-value	No. of cointegrating equation indicated
National (lags=2)	None *	34.061	0.0152	1 at the 5% level
	At most 1	9.834	0.2137	
	At most 2	0.751	0.3861	
Eastern (lags=3)	None*	33.280	0.0191	1 at the 5% level
	At most 1	11.639	0.1751	
	At most 2	0.222	0.6373	
Central (lags=2)	None *	41.462	0.0015	2 at the 5% level
	At most 1*	16.853	0.0310	
	At most 2	2.124	0.1450	
Western (lags=3)	None *	42.558	0.0010	2 at the 5% level
	At most 1*	20.701	0.0075	
	At most 2	2.3375	0.1263	

Note: * denotes rejection of the hypothesis at the 5% level, using MacKinnon-Haug-Michelis (1999) p-values.

Table 3 Likelihood ratio test of restrictions on the cointegrating vectors

Region	VAR lag	Coin. rank	Likelihood ratio statistics		
			[1,-1, 0]	[1, 0, -1]	[1, -1, 0] and [1, 0, -1]
National	2	1	6.07**	4.28	n.a.
Eastern	3	1	6.82**	1.90	n.a.
Central	2	2	6.92***	1.83	7.13**
Western	2	2	8.65***	3.70*	10.54***

Note: *, ** and *** denote significance at levels 10%, 5% and 1%.

Table 4 Cointegration tests with a structural break in 1978

Region	Hypothesized No. of CE(s)	Trace stat	Critical value at 5%	Critical value at 10%	No. of cointegration equations indicated
National	None	54.217	43.444	40.257	1 at 5% 2 at 10%
	At most 1	23.935	26.433	23.889	
	At most 2	9.220	12.846	11.049	
Eastern	None	53.258	43.444	40.257	1 at 5% 3 at 10%
	At most 1	25.995	26.433	23.889	
	At most 2	12.995	12.846	11.049	
Central	None	48.385	43.444	40.257	1 at 5% 1 at 10%
	At most 1	20.555	26.433	23.889	
Western	None	64.831	43.444	40.257	2 at 5% 3 at 10%
	At most 1	29.377	26.433	23.889	
	At most 2	12.020	12.846	11.049	

Notes: The Eviews programmes used to implement the method of Johansen *et al*. (2000) and to obtain the critical values for the tests are available on request.

Table 5 Unit root test by Perron's method, with a break in 1978

Table 5a	Log ratio of consumption to GDP			
	National	East	Central	West
μ	-0.221*** (-4.521)	-0.202*** (-4.522)	-0.325*** (-4.879)	-0.151*** (-3.793)
β	-0.003 (-1.580)	-0.005*** (-2.788)	-0.002 (-0.714)	0.0002 (0.140)
δ_1	0.158** (2.239)	0.084 (1.388)	0.294*** (2.929)	0.175** (2.472)
δ_2	-0.004 (-1.616)	-0.00001 (0.007)	-0.008** (-2.496)	-0.007*** (-2.777)
δ_3	-0.047 (-0.779)	-0.045 (-0.825)	-0.061 (-0.761)	-0.040 (0.690)
α	-0.569*** (-4.992)	-0.048*** (-4.934)	-0.903*** (-5.368)	-0.633** (-4.896)
Lag length	1	1	2	1
$\beta+\delta_2$	-0.0061***	-0.0047***	-0.0095***	-0.0067***
χ^2 statistic for H0: $\beta+\delta_2=0$	11.108	9.349	13.558	12.836
p-value	0.0009	0.0022	0.0002	0.0003

Table 5b	Log ratio of investment to GDP			
	National	East	Central	West
μ	-1.088*** (-4.823)	-0.96*** (-4.098)	-0.82*** (-3.895)	-1.106*** (-5.554)
β	0.005 (1.310)	0.002 (0.615)	0.0004 (0.077)	0.017*** (3.275)
δ_1	-0.231 (-1.334)	-0.177 (-1.024)	-0.215 (-1.030)	-0.201 (-1.131)
δ_2	0.010* (1.721)	0.012* (1.855)	0.010 (1.493)	-0.003 (-0.466)
δ_3	-0.046 (-0.310)	-0.059 (-0.407)	-0.114 (0.627)	0.092 (0.576)
α	-0.610*** (-5.085)	-0.515* (-4.175)	-0.482** (-4.348)	-0.679*** (-6.171)
Lag length	1	1	1	1
$\beta+\delta_2$	0.0151***	0.0139***	0.0107**	0.0139***
χ^2 statistic for H0: $\beta+\delta_2=0$	9.774	7.287	4.073	9.798
p-value	0.0018	0.0069	0.0436	0.0017

Note: Values in brackets are t statistics. *, ** and *** denote significance at levels 10%, 5% and 1%. The critical values for rejecting the null hypothesis of a unit root (α=0) from Table VIB (Perron, 1989) at λ=27/53=0.5 are -3.96, -4.24, and -4.90, respectively, for these significance levels.