Time Series Tests of Convergence of Chinese Regional and Provincial Economic Performance

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Abstract

This paper uses multivariate time series methods to investigate convergence of Chinese real GDP per capita at regional and provincial levels over the period 1952 – 2001. We reject convergence across regions. However, we find evidence of common trends among the provincial real incomes within Eastern and Central regions respectively while the cointegration relationships within the Western region are not sufficient to impose a common trend. We conclude that, contrary to the announced development policy of the Chinese government, the regions of China do not share a common development path.

JEL Classification: C23; O11; O41; O53

Key words: China; economic growth; regional convergence; cointegration, structural breaks

1. Introduction

A large amount of literature has been dedicated to investigating the convergence of China's regional per capita incomes. Most of these studies, such as Chen and Fleisher (1996), Jian, Sachs and Warner (1996), Li, Liu and Rebelo (1998), apply cross-sectional techniques. Convergence is typically defined operationally as a negative correlation between the initial level of income for a region and its subsequent economic growth rate. The general findings are as follows. There is little evidence in favor of an absolute convergence between 1952-1977. There appears to be absolute convergence during 1978-1989 while there is a divergence since 1990. However, there is possibly still a convergence since 1990 conditional on variables such as human capital investment and openness (Ghatak and Li, 2002).

However, the use of cross-section methods to investigate convergence has been exposed to some criticism. Quah (1993) argues that the investigation of dynamic behaviour via cross-sectional evidence risks inferential errors reminiscent of Galton's fallacy; a negative regression coefficient for initial levels in a cross-section study may in fact be perfectly consistent with the absence of convergence. Bernard and Durlauf (1996) show how cross-sectional methods can erroneously indicate convergence amongst a group of economies that in fact possess a variety of long run steady states. Islam (2003) provides a recent survey of the literature in which these arguments are embedded.

Carlino and Mills (1993) adopt a univariate time-series perspective for their investigation of regional convergence in the USA. The series of interest are those for regional per capita incomes *relative* to the national level and they define "stochastic convergence" as a lack of persistence of shocks, i.e. an absence of unit roots, in the individual series. This perspective permits regions to converge on a steady state in which individual regions may display permanent, i.e. equilibrium, differentials in relative per capita income. Zhang, Liu and Yao (2001) apply a similar approach and report evidence of such stochastic convergence for the Eastern and Western regions of China.

Our approach here is similar to Tsionas (2001). We follow Bernard and Durlauf (1995) who define convergence in terms of the long-run forecasts of incomes across economies and develop a multivariate testing apparatus founded in cointegration theory. Using their approach, we test for convergence across regions during 1952-2001 and explore whether convergence tendencies differ between the two sub-periods, 1952-1977 and 1978-2001, that precede and follow the implementation of significant reforms since 1978. In addition to this regional analysis, we investigate income convergence at provincial level within each of the regions.

The rest of the paper is organised as follows. Section 2 defines the concept of convergence in the context of cointegration analysis; section 3 outlines the policy context and the design of our empirical investigation; section 4 presents the statistical results and section 5 concludes.

2. Definition of Convergence in the Context of Cointegration Analysis

Following Bernard and Durlauf (1995, 1996) we express the concept of convergence in terms of the properties of time-series drawn from the economies under investigation. We define "convergence to equality" as the equality of the long-run forecasts of per capita income, y. With I_t, the information available at time t, and h as the forecast horizon, a conclusion of convergence to equality amongst a set of n economies requires

$$\underset{h\to\infty}{Lim}\, E\left(y_{1,t+h} - y_{p,t+h} \mid I_t\right) = 0 \quad \forall p \neq 1 \qquad \textbf{1.}$$

The economies, which may be separate countries or regional sub-divisions of an aggregate economy, are indexed by $p = 1,2,\ldots n$.

If economies do not converge in the sense of equality of the long-run forecasts of per capita income, they may still show "convergence to a common trend". This concept proposes that there exist a steady state pattern of relative incomes amongst the economies. Shocks to individual economies may disturb this pattern temporarily but such disturbance does not persist; the long run forecast is that the economies will converge upon their steady state pattern of relative incomes. Consequently the long run tendency is for per capita income in the several economies to rise or fall together, i.e. to share a common trend.

With $Y_t = \left(y_{1t} \quad y_{2t} \quad \ldots \quad y_{nt}\right)^T$ as a column-vector of per capita incomes and $c_p{}^T = \left(1 \quad -c_{p2} \quad -c_{p3} \quad \cdots \quad -c_{pn}\right), \quad p = 2,3\cdots n$ as *n-1* row-vectors of constants normalised to unity on the first economy, convergence to a common trend can be defined as

$$\underset{h\to\infty}{Lim}\, E\left(c_p{}^T\; Y_{t+h} \mid I_t\right) = 0 \quad \forall p \neq 1 \qquad \textbf{2.}$$

These *n-1* equations define the relativities amongst the *n* economies, leaving only one degree of freedom for a common trend.

The unit root and cointegration literature provides tools for testing the applicability of the above definitions. In this literature individual time series are described in terms of their autoregressive specification. If the lag polynomial that characterises the autoregressive behaviour of a series possesses a unit root then the series has zero probability of converging to a steady state so will tend to display local and/or persistent trends. Cointegration exists when combinations of series, for example $\left\{y_{1,t+k} - y_{p,t+k}\right\}$ in equations 1 or $\left\{c_p{}^T Y_{t+k}\right\}$ in equations 2, show evidence of convergence to a stationary steady state even though the individual series $\left\{y_{pt}\right\}$ are trended.

If a group of *n* individual series possess *n* cointegrating combinations then the stationarity of each and every one of these combinations provides sufficient restriction on the freedom of movement of the series as to guarantee that they jointly possess a static long run equilibrium. Since this situation then denies the possibility of trend behaviour in the individual series, it is not relevant for our discussion of whether a set of *developing* economies have convergent growth paths. However, if the n series possess *n-1* cointegrating combinations, for example as indicated in equations 2, then they are able to share a long-run growth trend and display short-run dynamics that imply convergence towards this.

This case of *n-1* cointegrating restrictions implies fixed relativities in the long run which, normalising on the first economy, offers an alternative characterisation of convergence to a common trend[1]. Using scalars k_p to indicate the relativities, we have

$$\underset{h \to \infty}{Lim}\, E\left(y_{1,t+h} - k_p y_{p.t+h} \mid I_t\right) = 0 \quad \forall p \neq 1 \qquad \textbf{3.}$$

Equations 3 state that with the passage of time the economies achieve fixed relativities in per capita income so that their proportional growth rates of per capita income are expected to become identical. Equations 1 can be seen as a further (testable) restriction of equations 3, in which the long run tendency is to share not only a common growth rate but also a common, growing, level of per capita income.

Obviously, the possibility exists that per capita incomes for a group of growing economies share fewer than *n-1* cointegrating restrictions. Particular examples could include situations such as m of the n economies sharing *m-1* restrictions that constrain them to a common trend with the remaining *n-m* economies sharing *n-m-1* restrictions that imply a different common trend. Here we might say that there is convergence *within* sub-groups but not *between* subgroups.

Tests of the above definitions of convergence can be developed within standard analysis of the cointegration properties of a column-vector of time series of per capita incomes, $Y_t = \left(y_{1t} \quad y_{2t} \quad \cdots \quad y_{nt}\right)^T$ - see for example Hamilton (1994).

[1]See appendix 1 for the derivation of equations 3.

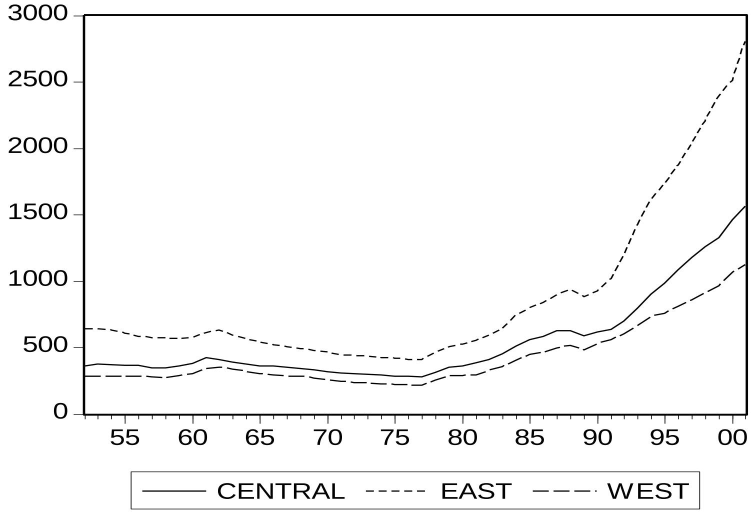
3. Motivation and Design of the Empirical Investigation

Throughout the period under study, the Chinese government has been implementing regional or unbalanced development strategies, albeit with an aim of achieving eventual equality across regions. The Chinese First Five-Year Plan (1953-1957) intended to correct a sub-optimal spatial distribution of industries by devoting most investments to inland (i.e. central and western) regions in preference to the richer coastal (eastern) one. This development strategy complemented the military considerations of that time. Projects in inland regions received additional funding via so-called Third Front strategy beginning in the mid-1960s while investment funded by the central state plan in the coastal region decreased. This development strategy could be said to address issues of equity at the expense of immediate efficiency in the distribution of support for investment projects.

Since 1978, this policy stance has been reversed. The geographical advantage of the coastal region has been used to integrate China with the outside world. Four special economic zones in the poor provinces of the eastern region, Guangdong and Fujian, were created in mid-1979, and further 14 coastal cities were opened in 1984. These zones and cities were granted considerable autonomy, preferential tax treatment and relatively large amounts of resources from the central government. The objective was to construct a more efficient policy that would promote growth in the coastal region with the idea that there would be spillover effects to the interior. To increase the opportunity for spill-over, reforms were implemented in the rural areas, where peasants were given autonomy for decision-making. A more balanced regional development strategy was not adopted until the late 1990s when Long-term Prospects for 2010 and the Western Development Strategy were drafted.

Figure 1 presents a time plot of the series of regional average incomes. Average income in all three regions stagnates and does not increase until the late 1970s. Over the whole period under study, 1952-2001, average income in the eastern region has been persistently higher than that of the central and western regions. The growth rate of income in the eastern region since 1978, indicated by the slope of the curves, is again higher than that of the other two regions. The time plot does not suggest convergence; in the next section we consider this question more formally.

Figure 1. Chinese regional incomes 1952-2001



The data used in the empirical exercise are annual provincial real GDP per capita over 1952-2001. Throughout the period under study, our data is based on 28 provinces; Tibet is dropped due to incompleteness of data, Hainan is combined with Guangdong and Chongqin is combined with Sichuan.

First, we investigate the question concerning convergence across Chinese regions during 1952-2001. There are three regions in China, namely, Eastern, Central and Western. We construct real GDP per capita for each region from GDP and population data for the provinces that make up each region. We look at the properties of the three regional time series, and test the convergence hypothesis. We further split the whole period into two sub-periods, 1952-1977 and 1978-2001, to see the effect of reforms since 1978.

Secondly, we examine convergence *within* regions over 1952-2001 using provincial real GDP per capita directly. We include 11 provinces in the eastern region, nine provinces in the central region and eight provinces in the western region.

4. Results of the Cointegration Analyses

4.1 Convergence between regions

We would like to formally investigate the apparent lack of regional convergence noted in the previous section by cointegration analysis. We consider both the full sample period, 1952-2001, and also the separate sub-periods prior to and following the change of policy stance noted above.

As is usual in convergence studies, we work with the natural logarithm of the real GDP per capita. We first test for stationarity and the order of integration in each of the three regional series of real GDP per capita. We use the Augmented Dickey-Fuller (ADF) test with a first order lag of the lagged differences, augmenting the autoregressive specification with intercept and trend[2].

Table 1 presents the results of the Augmented Dickey-Fuller tests. None of the three regions reject the null hypothesis of a unit root in level terms of income. Disregarding any possibility of structural breaks, the indication is that these series in level terms are individually non-stationary around linear time trends. The eastern and western regions reject the null hypothesis of a second unit root at a 5% level of significance and the central region rejects this I(2) possibility almost as strongly. We conclude that the three series are integrated of order one, I(1). The faster growth of the second half of the data set is therefore implicitly assigned to the year on year integration of the linear trend components.

Table : Unit root tests on log real per capita GDP during 1952-2001

	ADF (levels)	ADF (first differences)
Eastern region	-0.695	-4.015
Central region	-0.863	-3.358
Western region	-0.811	-3.647

Note: MacKinnon critical values for rejection of hypothesis of a unit root in level terms: -4.158 at 1%, -3.505 at 5% and -3.182 at 10% level of significance. MacKinnon critical values for rejection of hypothesis of a unit root in the first differences: -4.163 at 1%, -3.507 at 5% and -3.183 at 10% level of significance.

[2] The lag length in the ADF test model was chosen so as to remove serial correlation in the residuals.

Although the three series are non-stationary in level terms, they may possess cointegrated linear combinations that are stationary. To test if the series are cointegrated, we model them jointly as a vector error correction model (VECM) and apply Johansen's procedure to estimate the number of stochastic trends. Since the ADF test regressions for Table 1 incorporated linear trends, we specify the VECM with a deterministic trend in the series. We additionally specify an intercept but no trend in the cointegration equations so that our null hypothesis is of convergence upon a pattern of relative incomes that is constant but might not involve equality across the regions. We use the sequential modified LR test statistic at 5% level of significance for selecting VAR lag order, with the additional requirement that diagnostic statistics suggest that the model's errors are multivariate normal. In this case of 1952-2001 the LR test indicates that the appropriate lag order is 2 in the VAR and the probability of rejecting the null of multivariate normality of residuals by Cholesky Jarque-Bera normality test is 0.226.

Table : Johansen cointegration test for the regional incomes, 1952-2001

Hypothesized No. of CE(s)	Eigenvalue	Trace Statistic	5 Percent Critical Value	1 Percent Critical Value
None	0.202947	17.06739	29.68	35.65
At most 1	0.112721	6.179330	15.41	20.04
At most 2	0.009099	0.438758	3.76	6.65

*(**) denotes rejection of the hypothesis at the 5%(1%) level
Trace test indicates no cointegration at both 5% and 1% levels

Table 2 reports the results of the Johansen procedure for the period 1952-2001. As the value of the likelihood ratio is less than the critical value at 5% level of significance, we cannot reject the null hypothesis that the rank of the cointegrating matrix is zero. This implies three separate stochastic trends driving the series, without any long-run relationship between them in the system. We can conclude that the statistic shows no tendency for the three regions to converge in terms of real per capita income. The strategies of favoritism to the inland regions during the 1950s and 1960s and preferential treatment to the coastal regions in later years do not rectify the disparity of real incomes across regions as intended by the Chinese government. Given the government's motivation towards shared growth as a means towards social and economic stability, this could be seen as a failure of development policy to deliver the desired outcomes.

Since there was a considerable change in the regional development strategy in 1978, we would like to see whether there is any difference in the tendency to converge before and after the reforms were implemented. We split the whole period into two sub-periods, 1952-1977 and 1978-2001. Before we can carry out the cointegration analysis, we need to confirm that the three series are individually I(1) during each sub-period. When testing the stationarity of the series during 1952-1977 and 1978-2001, we again specify the ADF test regression with an intercept and trend and select lag length so as to achieve stationarity in the residuals.

Table : Unit root tests for the sub-periods

	ADF (1952-1977)			ADF (1978-1998)	
	Levels	First differences	Second differences	Levels	First differences
East	-2.515	-3.145	-4.282	-2.182	-3.869
Central	-1.725	-2.787	-3.787	-2.331	-3.2414
West	-1.82	-2.452	-3.381	-3.374	-5.162

Note: During 1952-1977, MacKinon critical values for rejection of null hypothesis of a unit root in level terms: -4.394 at 1%, -3.6118 at 5% and –3.242 at 10% level of significance. MacKinon critical values for rejection of null hypothesis of a unit root in first differences: -4.447 at 1%, -3.787 at 5% and –3.247 at 10%. MacKinon critical values for rejection of null hypothesis of a unit root in second differences: -4.44 at 1%, -3.63 at 5% and –3.253 at 10%. During 1978-2001, MacKinnon critical values for rejection of hypothesis of a unit root in level terms and first differences: -4.394 at 1%, -3.612 at 5% and -3.2418 at 10% level of significance.

In the second sub-period the ADF test procedure confirms for all three regions the I(1) behaviour previously diagnosed for the complete period. In the earlier sub-period however, there is a suggestion that the series are integrated of order I(2), which is counter to expectations. Since it is well-known that a stationary series which experiences structural breaks in its deterministic component may be mis-diagnosed by the ADF procedure as being I(1), we consider the possibility that the diagnosis of I(2) behaviour is in fact the result of I(1) processes experiencing structural breaks.

The apparent I(2) behaviour of the series of regional per capita incomes during 1952-1977 may have been due to a structural break caused by the Great Leap Forward during 1958-1961. Chow (2002, Chapter 8) estimates that in the absence of the “Great

Leap Forward" output and consumption per worker by 1992 might have been twice as large as that actually observed. He further suggests that the general economic disruption caused by this episode may have exceeded that of the later "Cultural Revolution". In 1958, the Great Leap Forward was launched with the purpose of increasing China's output dramatically and developing its economy rapidly. In rural areas farmers were organised into communes and required to work cooperatively in pursuit of overly ambitious output targets. In the cities, people were asked, for example, to build furnaces in their backyards to produce iron and steel. To meet output targets for raw metal, finished products were put into the furnaces. The end result was an economic disaster. Food production was greatly reduced and over 25 million people died of famine from 1958-1962.

In order to model the effect of the Great Leap Forward we apply the method of Perron (1989). This involves a test regression in which the null and alternative hyptheses each permit the possibility of a single structural break at a date selected by the investigator; we choose 1961. We assume that the events of 1958-1961 change not only the intercepts but also any deterministic trends in the series.

The strategy is to nest null and alternative hypotheses within a single model.

$$\Delta y_t = \mu + \beta t + \delta_1 DU_t + \delta_2 DT_t + \delta_3 D_t + \alpha y_{t-1} + \sum_{i=1}^{i=k} \gamma_i \Delta y_{t-i} + \varepsilon_t \qquad 4.$$

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

$D_t = 1$ for $t= T_B +1$, i.e. 1962, otherwise $D_t = 0$.

$DU_t = 1$ for $t > T_B$, i.e., from 1962, otherwise $DU_t = 0$.

$DT_t = t$ for $t > T_B$, i.e., from 1962, otherwise $DT_t = 0$.

The null hypothesis is that the series is generated by a unit root process which has no deterministic trend but experiences a jump and a change of drift rate at the break and requires $\beta=0$, $\delta_2=0$ ("no time trend") and $\alpha=0$ ("unit root"). The alternative hypothesis is of stationary fluctuations about a deterministic trend that changes its intercept and slope at the break date: $\alpha<0$ ("error correction"), $\delta_1<>0$ ("change of intercept"), $\delta_2<>0$ ("change of slope").

Although the null hypothesis is a point in 3-D parameter space it is common practice to make a judgement on the basis of the "Dickey-Fuller t-statistic" for the coefficient α. Perron (1989) considers the asymptotic behaviour under null and alternative hypotheses of a number of test statistics including the familiar ADF t-statistic. He presents tables of percentiles for the asymptotic distribution of the test statistics. These critical values are more extreme than for the standard ADF test and the disparity is larger as the break point is closer to the centre of the sample.

Due to the small number of observations, we drop the lagged dependent variable from the right-hand side of the equation 4. The following table reports the estimated equation for each region.

Region	μ	β	δ_1	δ_2	δ_3	α
East	-0.342 (1.015)	0.009 (0.003)	-0.054 (0.043)	-0.004 (0.003)	0.082 (0.02)	0.046 (0.157)
Central	-0.85 (1.58)	0.01 (0.004)	-0.08 (0.107)	-0.005 (0.008)	0.001 (0.033)	0.139 (0.267)
west	-0.193 (1.43)	0.012 (0.004)	-0.052 (0.156)	-0.008 (0.01)	0.08 (0.034)	0.027 (0.255)

Note: values in the brackets are standard errors.

The above results suggests that we can accept the existence of a unit root ($\alpha=0$) in the three region: the three series are integrated of order 1. Meanwhile the signs of the time trend and changing trend are consistent with the actual observation. Before 1962, the regional incomes increase about 1% per annum. But since 1962, the three regional incomes increases by 0.46%, 0.6% and 0.33% per annum respectively.

To see if the series are cointegrated during 1952-1977, we use Johansen's test without consideration of structural breaks. This constitutes an assumption that the effects on each region of the Great Leap Forward were sufficiently homogenous and contemporaneous as to imply "co-breaking" in the sense of Clements and Hendry (1999, Chapter 9), which may warrant further research. We assume that a test model with no deterministic trend in the data and with an intercept but no trend in the cointegrating equation is appropriate. The sequential modified LR test statistic indicates that the lag order should be 2 and the prob. Value for the test of multivariate normality of the residuals is 0.1903. Table 4 presents the results of this Johansen test for the subperiod, 1952-1977.

Table : Johansen cointegration test for the regional incomes, 1952-1977

Hypothesized No. of CE(s)	Eigenvalue	Trace Statistic	5 Percent Critical Value	1 Percent Critical Value
None *	0.659714	40.41629	34.91	41.07
At most 1	0.381341	14.54501	19.96	24.60
At most 2	0.118245	3.020184	9.24	12.97

*(**) denotes rejection of the hypothesis at the 5%(1%) level

In table 4, the likelihood ratio test at 5% level of significance indicates that there exists one cointegration relation between the series, and thus by implication two common trends. The system can diverge without bound in two independent directions and it is stable in only one direction. We conclude that whilst there is no tendency to converge in incomes during 1952-1977, there is a mechanism that to some extent constrains the deviation between regions. It seems that the development strategy in this pre-reform period could have partly constrained divergence between regions.

For the period 1978-2001, we specify the Johansen test model with a linear trend in data and an intercept but no trend in the cointegrating equation. Both the sequential modified LR test statistic and Akaike Information Criteria indicate that the lag order should be 1 and the Jarque-Bera test for multivariate normality of the residuals supports this decision, giving a prob. Value of 0.317. Table 5 reports the Johansen test on the regional incomes over 1978-2001.

Table : Johansen cointegration test for the regional incomes, 1978-2001

Hypothesized No. of CE(s)	Eigenvalue	Trace Statistic	5 Percent Critical Value	1 Percent Critical Value
None	0.442688	23.00545	29.68	35.65
At most 1	0.299332	8.974338	15.41	20.04
At most 2	0.018045	0.437031	3.76	6.65

*(**) denotes rejection of the hypothesis at the 5%(1%) level

As reported in Table 5, the likelihood ratio test at 5% and 1% levels of significance indicates that there does not exist any cointegration relation between the series. During 1978-2001, the series tend to deviate from each other, the rich eastern region having grown at a faster pace while poor central and western regions having grown at slower rates. It seems that 24 years included in this period is not long enough for the regional unbalanced development strategy implemented between late 1970s and 1990s to achieve the spillover effects of openness.

In summary, whilst the policies prior to 1978 may have limited the extent of divergence between regions, they were not sufficient to establish full convergence. In the period following 1978 the regions appear to be following separate growth paths with no evidence at all of convergence. This complete absence of convergence is also the conclusion that emerges from considering the data period as a whole.

4.2 convergence within regions

Although there is no evidence of convergence across regions during the period under study, we wonder if there is convergence within regions. We now work directly with data on natural logarithms of the provincial real GDP per capita during 1952-2001. First of all we carry out the unit root test on the series of provincial incomes to see if they are integrated of the same order within regions. We include an intercept and trend in the standard ADF test model. The residuals from the test models have been checked for stationarity. The results, reported in Table 6, indicate that all the income series are integrated of order I(1), so that Johansen's method may be applied to groups of these series in order to test for convergence.

Table : Unit root tests on log real GDP per capita, 1952-2001

Eastern region			Central region			Western region		
	ADF tests			ADF tests			ADF tests	
	x_t	Δx_t		x_t	Δx_t		x_t	Δx_t
Beijing	-2.79	-4.276	Shanxi	-1.274	-4.419	Sichuan	-0.967	-3.782
Tianjin	-0.337	-4.178	Inner Mongolia	-0.95	-4.382	Guizhou	-1.277	-3.852
Heibei	-0.241	-4.314	Jilin	-0.533	-4.61	Yunnan	-0.772	-3.84
Liao Ning	-0.79	-4.254	Heilongjiang	-0.168	-4.652	Shaanxi	-0.44	-4.442
Shang Hai	-1.124	-4.912	Anhui	-1.35	-3.46	Gansu	-0.328	-4.293
Jiangsu	-0.925	-3.696	Jiangxi	-0.584	-4.44	Qinghai	-0.83	-4.429
Zhejiang	-1.003	-4.009	Henan	-0.937	-3.202	Ningxia	-0.74	-5.553
Fujian	-0.606	-3.499	Hubei	-0.538	-3.787	Xinjiang	-0.963	-3.249
Shandong	-1.055	-3.504	Hunan	-0.547	-4.199			
Guangdong	-0.366	-4.158						
Guangxi	-0.854	-3.815						

Note: MacKinnon critical values for rejection of hypothesis of a unit root in level terms: -4.1584 at 1%, -3.5045 at 5% and -3.1816 at 10% level of significance. MacKinnon critical values for rejection of hypothesis of a unit root in first differences: -4.163 at 1%, -3.5066 at 5% and -3.182 at 10% level of significance.

Table 7 presents Johansen's procedure for estimating the dimensions of cointegration amongst incomes in the eleven provinces of the eastern region. The sequential modified LR test statistic for VAR lag order selection at 5% level of significance indicates that the lag order is 2. Ten cointegrating vectors are found by the trace statistic, using 5% critical values. The eleven provinces therefore share a single common trend.

Table : Cointegration within the Eastern region, 1954 - 2001

Hypothesized No. of CE(s)	Eigenvalue	Trace Statistic	5 Percent Critical Value	1 Percent Critical Value
None **	0.900593	552.7239	277.71	293.44
At most 1 **	0.884398	441.9141	233.13	247.18
At most 2 **	0.833324	338.3492	192.89	204.95
At most 3 **	0.757764	252.3476	156.00	168.36
At most 4 **	0.714084	184.2910	124.24	133.57
At most 5 **	0.605465	124.1923	94.15	103.18
At most 6 **	0.464501	79.55004	68.52	76.07
At most 7 *	0.298907	49.57135	47.21	54.46
At most 8 *	0.278058	32.52588	29.68	35.65
At most 9 *	0.248666	16.88695	15.41	20.04
At most 10	0.063781	3.163499	3.76	6.65

*(**) denotes rejection of the hypothesis at the 5%(1%) level

For the central region, a test model with a linear trend in data and with an intercept and no trend in the cointegrating equations is specified on the basis of the ADF test in Table 6. The sequential modified LR test statistic and Akaike Information Criteria for selecting VAR lag order indicate that the lag of VAR should be 3. Table 8 reports the Johansen test results. The Johansen test at 5% level of significance indicates that there are eight cointegrating vectors in the system of nine provinces in the central region, leaving a single common trend.

Table : Cointegration within the Central region, 1955 - 2001

Hypothesized No. of CE(s)	Eigenvalue	Trace Statistic	5 Percent Critical Value	1 Percent Critical Value
None **	0.942861	458.9962	192.89	204.95
At most 1 **	0.845385	324.4698	156.00	168.36
At most 2 **	0.814310	236.7294	124.24	133.57
At most 3 **	0.709092	157.5966	94.15	103.18
At most 4 **	0.555589	99.56344	68.52	76.07
At most 5 **	0.420035	61.44615	47.21	54.46
At most 6 **	0.346752	35.84117	29.68	35.65
At most 7 *	0.281325	15.82867	15.41	20.04
At most 8	0.006413	0.302379	3.76	6.65

*(**) denotes rejection of the hypothesis at the 5%(1%) level

For both eastern and central regions, there is a common stochastic trend among the provincial real incomes. This is a meaningful relationship among them i.e. the provinces within each region are apparently converging to a long-run equilibrium pattern of relative incomes. This finding for the eastern region is consistent with that achieved by the different approach in Zhang, Liu and Yao (2001). The coastal development strategy between late 1970s and 1980s has led to faster growth in the relatively poor provinces of Guangdong and Fujian in the eastern region, which eventually caught up with the rich provinces in the region. Meanwhile, the original industrial provinces with a high proportion of state-owned enterprises in the central region like Heilongjian and Jilin, which have not directly benefited from the policy of openness or rural reforms grew at a slower pace. The agricultural provinces in the central region grew at a relatively faster pace due to rural reforms. The provinces converged to the mean income of lower level in the central region.

For the western region, a test model with a linear trend in data and with an intercept and no trend in the cointegrating equations is selected on the basis of ADF test in Table 6. The sequential modified LR test statistic for VAR lag order selection indicates that the lag of VAR should be 2. Table 9 reports the Johansen test results. The Johansen test indicates that there are four cointegrating vectors in the system of eight provinces in the western region. The variation of incomes among provinces in the western region is constrained but the constraints are not sufficient to induce a common trend. The incomes do not converge in the western region. Our findings about the central and western regions are contrary to those made by Zhang, Liu and Yao (2001).

Table : Cointegration within the Western region, 1954 - 2001

Hypothesized No. of CE(s)	Eigenvalue	Trace Statistic	5 Percent Critical Value	1 Percent Critical Value
None **	0.775180	232.8940	156.00	168.36
At most 1 **	0.626177	161.2562	124.24	133.57
At most 2 **	0.582513	114.0256	94.15	103.18
At most 3 *	0.439391	72.09749	68.52	76.07
At most 4	0.368062	44.31839	47.21	54.46
At most 5	0.264683	22.28809	29.68	35.65
At most 6	0.144992	7.530297	15.41	20.04
At most 7	0.000236	0.011341	3.76	6.65

*(**) denotes rejection of the hypothesis at the 5%(1%) level

Conclusion

This paper makes use of the half a century data on Chinese provincial real GDP to investigate the issue of convergence by cointegration analysis. We do not find evidence of convergence *between* Chinese regions. However, convergence is evident *within* the eastern and central regions while there is no evidence that the provinces in the western region are converging. The implication of these findings is that the three regions do not share a common growth path. The Chinese government has rightly adjusted their priority to public investment in the western region in current Five-Year Plan (2001-2005) to facilitate the spillover effect intended in the earlier unbalanced development strategies.

This exercise enriches the methodology of empirical study on Chinese convergence. Future directions for extending the investigation include a data-based confirmation of the significance of changes of regime and an investigation of the propagation pattern for spillover effects.

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Appendix 1

Here we derive equations 3.

We begin with equations 2, $\underset{h\to\infty}{Lim}\, E\left(c_p{}^T\, Y_{t+h} \mid I_t\right) = 0 \quad \forall p \neq 1$, expressing the detail as a partitioned matrix equation:

$$\underset{h\to\infty}{Lim}\, E\left(\begin{bmatrix} \underset{(n-1)x1}{\iota} & \underset{(n-1)x(n-1)}{-C} \end{bmatrix} \begin{bmatrix} y_{1,t+h} \\ y_{2,t+h} \\ \vdots \\ y_{n,t+h} \end{bmatrix} \mid I_t \right) = \underset{(n-1)x1}{0},$$

in which ι is a vector of units and the elements of C are the c_{pi} defined above for equations 2.

We assume that the $c_p{}^T = \left(1 \quad -c_{p2} \quad -c_{p3} \quad \cdots \quad -c_{pn}\right)$ which are the rows of C are linearly independent so that C has an inverse. Multiplying through by this inverse gives

$$\underset{h\to\infty}{Lim}\, E\left(\begin{bmatrix} \underset{(n-1)x1}{C^{-1}\iota} & \underset{(n-1)x(n-1)}{-\mathrm{I}} \end{bmatrix} \begin{bmatrix} y_{1,t+h} \\ y_{2,t+h} \\ \vdots \\ y_{n,t+h} \end{bmatrix} \mid I_t \right) = \underset{(n-1)x1}{0}$$

and detailing the *(n-1)x1* column vector $C^{-1}\iota$ as $\left(1/k_2 \quad 1/k_3 \quad \cdots \quad 1/k_n\right)^T$ permits

$$\underset{h\to\infty}{Lim}\, E\left(\left[y_{1,t+h} \begin{bmatrix} 1/k_2 \\ 1/k_3 \\ \vdots \\ 1/k_n \end{bmatrix} - \begin{bmatrix} y_{2,t+h} \\ y_{3,t+h} \\ \vdots \\ y_{n,t+h} \end{bmatrix} \right] \mid I_t \right) = \underset{(n-1)x1}{0}$$

The individual equations of this block are presented in the text as equations 3.