
Rising regional inequality in China: Policy regimes and structural changes*

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Abstract. Regional inequality is severe in China since regional development is uneven due to various initial conditions and government policies. We employ unit root tests allowing for structural breaks to alternative inequality measures from 1952 to 2000. Empirical results indicate that (1) the regional inequality is trend stationary with structural breaks rather than following a random walk. Thus, ignoring structural changes might leads to incorrect inference and misleading policy implications; (2) the break points are associated with episodic events in Chinese economic history such as the Cultural Revolution and market reforms. This implies that the policies had a long-lasting and fundamental effect on the inequality.

JEL classification: C22, O15, R58

Key words: Structural break, unit root, regional inequality, China

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1 Introduction

Rising regional inequality in China has attracted attention from many groups of people. The regional inequality is severe since regional development is uneven due to various initial conditions (e.g. different geographies) and government policies. The Gini coefficient of income increased from 0.3 in 1981 to more than 0.4 in 2005. Moreover, the uneven development between coastal and inland regions became aggravated. The ratio of per capita GDP in Shanghai to that in Guizhou province in 2005 is 9.9 (Asian Development Bank 2006). Lin (2005) points out that the root cause for the high inequality in China is not because the rich people are too rich, but because the poor people are too poor. The critical aspect for closing the income gap relies on increasing the income levels of poor people and helping them out of the poverty trap. With the expanding income gap, social instability has surfaced as a consequence. Income inequality, as documented in Song and Appleton (2006), lowers the subjective perception on satisfaction of people's well being. Policymakers in China have come to realise the seriousness of inequality and measures must be taken to improve the economic and social conditions for the poor.

There are numerous studies identifying the determinants of inequality in China. Kanbur and Zhang (2005) find that openness is positively related to inequality because trade liberalisation and foreign direct investment (FDI) change the internal comparative advantage and location choice of firms as provinces closer to ports attract more trade and investment from abroad (Xing and Zhang 2004). More fiscal decentralisation is associated with greater inequality, as shown in Tsui (1991) and Kanbur and Zhang (2005). However, there is little discussion regarding the empirical properties of the inequality, an issue which is useful for policy analysis in itself.

In this paper, we examine the time series properties of inequality in China. In particular, we want to show whether the evolution of inequality is non-stationary, trend stationary or trend stationary with structural breaks. If the inequality is non-stationary, then changes in the inequality are permanent. It questions the efficacy of redistribution policies since the impacts of these policies can be offset by other shocks on inequality. However, if the inequality is trend stationary, it implies that a government-initiated structural reform will have a long-lasting effect on inequality since shocks on the inequality cannot alter the underlying trend¹. However, if the inequality is actually trend stationary with structural breaks in level and/or slope, standard unit root tests have little power to reject the null of unit root and hence inferences from these tests might be misleading.

By examining the direction and timing of changes in the evolution of inequality, this paper contributes in two aspects. First, it enables us to track down the response of inequality to the various policies adopted in modern China, informing policymakers about future policies geared to close the income gap. Second, it contributes to the empirical literature of using time series technique to explore inequality. Hayes et al. (1990) look into the income inequality in the United State and find that the size distribution of income in the U.S. follows a random walk.² Employing the aggregate and provincial data (except New Brunswick) of Canada, Wang and Ogwang (2004) show that the Gini coefficients are non-stationary. However, the statistical properties

¹ The trend function explains the movement of inequality within a sample but does not extrapolate for forecasting. For example, even if the trend is positive, it is unlikely that the Gini coefficient will go to one since structural change in the economy can occur in the future, which stabilise or reverse the trend of the income inequality.

² Raj and Slotte (1994) and Rothman (1998) reject the null hypothesis of unit root in the income inequality measures of the U.S. by employing unit root test with one structural break and higher power, respectively.

of inequality in China have been little studied.³

Our empirical results show that the conventional unit root tests without structural breaks cannot reject the unit root hypothesis. By contrast, unit root tests admitting the structural breaks illustrate that the alternative inequality measures for China are trend stationary with breaks rather than following a random walk. Moreover, the break points are associated with episodic events such as the Cultural Revolution and market reforms. This evidence suggests that policies can be effective to alleviate inequality by altering the long-run trend in China.

The rest of the paper is organised as follows. The next section provides a brief introduction to Chinese modern economic history. Data are discussed in the third section. The fourth section explains various unit root tests. The empirical results are reported in the fifth section. The last section provides a conclusion.

2 Historical background⁴

The People's Republic of China was established in 1949. After the economic recovery of 1950–52, the Chinese government started to focus on economic development to rebuild and industrialise the country through the First Five-year Plan from 1953 to 1957. This period was the central planning period, during which the government allocated most resources to heavy industries, controlled interest rates, trade and financial sectors, and expanded the share of production of state-owned enterprises (SOEs) in the economy. The average annual GDP growth was 9.1% during this period. In the Second Five-year Plan period (1958–1962), the authorities adopted a policy called the Great Leap Forward to speed up the pace of modernisation of the economy in order to catch up with industrial countries. However, the failure of this radical plan led the GDP to decline precipitously by 2.2% annually. From 1963 to 1965, the Chinese economy entered a period of adjustment and recovery, which was driven by a series of corrective measures. Unfortunately, the Cultural Revolution (1966–1976) obliterated these developments and the pace of economic growth slowed down again.

The following decades are called the reform period, and which can be divided into three different stages denoted as stage one (1978–1984), stage two (1985–1991) and stage three (1992–present). Deng Xiaoping came into power after the Third Plenary Session of the Eleventh Chinese Communist Party Congress in December 1978, and this marks the era reform period up to now. The first stage of the reform started by implementing the household responsibility system which linked remuneration to output for agricultural production in the late 1970s. As a result, the total agricultural growth was 42.6%, which solved the shortage of agricultural products and improved people's well-being. Furthermore, the government decided to establish four special economic zones (SEZs)⁵ to attract FDI.⁶

In the second stage, the government expanded the reforms to include SOEs in the urban areas. It delegated more decision-making power to SOEs and let them be

³ Researchers have employed time series techniques to look into the GDP at aggregate (Li 2000), regional (Zhang et al. 2001; Groenewold et al. 2007) and provincial levels (Smyth and Inder 2004) to examine income convergence in China.

⁴ Further details on the development of the Chinese economy can be found in Lin et al. (2003)

⁵ The SEZs were set up in the cities of Shenzhen, Zhuhai and Shantou in Guangdong province in 1980 and Xiamen in Fujian province in 1981.

⁶ However, FDI inflow did not increase until 1984. In the same year, the Chinese government decided to open another 14 coastal cities and Hainan island (which became a province in 1988). A detailed timeline of preferential policy in China can be found in Demurger et al. (2002).

responsible for profits and losses through contracting and leasing. Under the contract, SOEs were allowed to retain extra profits after fulfilling the contracted quota and have control over personnel decisions. In order to improve managerial incentives, the shareholding system⁷ and tax reform⁸ were introduced.

In early 1992, Deng Xiaoping's speech during his visit to southern China began the third stage of reform. It confirmed the role of the market economy and sped up the reform further through the opening of the market. According to Fung et al. (2004), from 1991 to 2000, the Chinese trade volume increased from U.S. \$53.4 billion to \$225.1 billion while FDI inflows increased dramatically from US\$4.4 billion to \$40.7.

In summary, many policies were introduced during the Chinese modern historical period, which affected the economy in various tremendous ways. Thus, it is difficult to choose one or two exogenous break points for our data based on our ex post knowledge of it. Moreover, the stage of economic development varies across cities, provinces and regions. Thus policy lags in some parts of the country may cause the break points in the data to deviate from the so-called historical turning points in Chinese economic history. Therefore, allowing endogenous structural breaks in our tests is appropriate and desirable.

3 Data description

The regional inequality measures used in this paper are constructed by Kanbur and Zhang (2005) using real per capita consumption at the provincial level (except Tibet and Hainan provinces). There are four measures of inequality covering the sample period from 1952 to 2000. There are two overall inequality measures, namely, the Gini and the decomposable generalised entropy (GE) class of inequality measure. The other two are urban-rural inequality and inland-coastal inequality GE measures. The data are listed in the appendix in Table A1.

The Gini coefficient is defined as the ratio of the average difference between all possible pairs of consumption and the total consumption of the economy, i.e.

$$G = \frac{\frac{1}{2n} \sum_{i=1}^n \sum_{j=1}^n |y_i - y_j|}{Y} \quad (1)$$

where n and Y are the total population and total consumption expenditure in the economy, respectively, and y_i and y_j are the consumptions of the i -th and j -th province, respectively.

The decomposable GE class of inequality measures are developed by Shorrocks (1980, 1984):

$$I(y) = \sum_{i=1}^n f(y_i) \log\left(\frac{\mu}{y_i}\right) \quad (2)$$

where y_i is i^{th} consumption expenditure measured as Chinese Yuan and y is a vector, μ is the total sample mean and $f(y_i)$ is the population share of y_i in the total population n . The GE measure is additively decomposable:

⁷ In July of 1984, the first shareholding company, Beijing Tianqiao department store (shareholding), was formed. In 1991, at the end of the second stage of reform, there were 709 SOEs restructured to the shareholding system.

⁸ SOEs paid taxes on their revenue to state and local governments instead of handling in profit. This policy aimed to loose the relation between SOE and local authorities.

$$I(y) = \sum_{g=1}^k w_g I_g + I(\mu_1 e_1, \dots, \mu_k e_k) \quad (3)$$

where $w_g = n_g/n$ represents the share of the g^{th} group's population in the total population. I_g is the inequality of the g^{th} group, μ_g is the mean of the g^{th} group, and e_g is a vector of 1's of length n_g . The first term is the within group inequality whereas the second term is the between group inequality. Kanbur and Zhang (2005) decompose the GE into within and between components for rural-urban⁹ and inland-coastal¹⁰ inequalities.

4 Methodology

To determine the time series properties of the inequality measures, we employ three unit roots tests without structural breaks and two with structural breaks. The tests without breaks serve as a comparison for the effect of adding endogenous breaks into the test procedure.

4.1 Unit root tests without structural breaks

First, the Augmented Dickey-Fuller (1981) (ADF) unit root test is applied to the natural logs of those four inequality measures with the following specification:

$$\Delta y_t = \alpha + \beta t + \mu y_{t-1} + \sum_{i=1}^k c_i \Delta y_{t-i} + \varepsilon_t \quad \text{where } t=1,2,\dots,T \quad (4)$$

where Δ is the first difference operator, y_t is one of the inequality measures, t is the time trend, and ε_t is an independently and identically distributed process with mean 0 and variance σ^2 . In order to ensure the error term is white noise, Dickey and Fuller (1981) suggest handling serial correlation using the lagged differenced terms. The null hypothesis is that $\mu = 0$, which implies there is a unit root in y_t . The alternative hypothesis is that $\mu < 0$ which implies y_t is stationary around a deterministic trend.

Second, a Phillips-Perron (1988) (PP) test is conducted. The PP test is also based on the similar equation as employed in the ADF test, but without the lagged differenced terms:

$$\Delta y_t = \alpha + \beta t + \mu y_{t-1} + \varepsilon_t \quad \text{where } t=1,2,\dots,T \quad (5)$$

While the ADF test corrects for higher-order serial correlation by adding lagged differenced terms as independent variables, the PP test makes a nonparametric correction to account for the autocorrelation. Monte Carlo studies suggest that the PP test generally has greater power than the ADF test (Banerjee et al. 1993).

The third is the Dickey-Fuller Generalised Least Squares (DF-GLS) test, proposed by Elliot et al. (1996), which is a modified version of the Dickey-Fuller test. Define

⁹ Set $K=2$. Kanbur and Zhang use real per capita consumption in rural and urban areas to compute within rural-urban inequality, the first term of equation (3). The difference between the overall inequality and the within rural-urban inequality is the between rural-urban inequality, i.e., rural-urban inequality in the Appendix Table A1.

¹⁰ Set $K=2$. The coastal group includes the provinces of Beijing, Liaoning, Tianjin, Hebei, Shandong, Jiangsu, Shanghai, Zhejiang, Fujian, Guangdong and Guangxi. Kanbur and Zhang use real per capita consumption in the coastal and inland provinces to compute within coastal-inland inequality, the first term of equation (3). The difference between the overall inequality and the within coastal-inland inequality is the between coastal-inland inequality, i.e. coastal-inland inequality in the Appendix Table A1.

$$\bar{y} = [y_1, (1-\alpha L)y_2, \dots, (1-\alpha L)y_T]' \quad (6)$$

$$\bar{z} = [z_1, (1-\alpha L)z_2, \dots, (1-\alpha L)z_T]' \quad (7)$$

where $\alpha = 1 + (\bar{c}/T)$, L is the lag operator, y_t is the original time series and z_t is $[1, t]'$. The DF-GLS test proceeds by first de-trending the series as

$$y_t^d = y_t - \tilde{\beta}_0 - \tilde{\beta}_1 t \quad (8)$$

where $\tilde{\beta}_0$ and $\tilde{\beta}_1$ are obtained by regressing \bar{y} on \bar{z} . In order to improve the power of the tests, Elliot et al. (1996) suggest the parameter \bar{c} takes the value -7 or -13.5 if the original data generation process (DGP) is assumed to have a constant or both constant and trend, respectively. Then, the DF-GLS test uses the following specification

$$\Delta y_t^d = \mu \Delta y_{t-1}^d + \sum_{i=1}^k \eta_i \Delta y_{t-i}^d + \varepsilon_t \quad \text{where } t = 1, 2, \dots, T \quad (9)$$

The DF-GLS test has the same null and alternative hypotheses as the traditional ADF test.

The empirical results for these three tests are summarised in Table 1. For these four inequality measures, the null hypothesis of unit root cannot be rejected even at a 10% significance level. It suggests that the inequality measures are integrated of order one. However, inspecting the time plots of the inequality measures in Figure 1, these measures may involve structural breaks. Therefore, we suspect the validity of the empirical results from these conventional tests because these models might be mis-specified.

<Table 1 about here>

<Figure 1 about here>

4.2 Unit root tests with structural breaks

Perron (1989) points out that the power to reject the null of unit root decreases when the stationary alternative is true, but a structural break (SB) is ignored. He proposes a modified ADF test to allow one exogenous SB for three possible models: a model with shift in intercept, a model with change in slope and a model with changes in both intercept and slope. In Zivot and Andrews (1992) (ZA), they argue that the break points should be viewed as being correlated with the data. Therefore, selecting the breaking points a priori based on an ex post knowledge of the data could lead to an over-rejection of the unit root hypothesis. In terms of inequality in China, it is difficult to pick up a few break points during the last fifty years since the Chinese economy experienced many fundamental changes. Thus, we follow ZA to treat the break as endogenous, which is determined within the model. The specification of the ZA model with only shift in intercept (ZA1) is

$$\Delta y_t = \alpha + \beta t + \gamma DI + \mu y_{t-1} + \sum_{i=1}^k c_i \Delta y_{t-i} + \varepsilon_t \quad \text{where } t = 1, \dots, T \quad (10)$$

The model with shift both in intercept and slope (ZA2) is

$$\Delta y_t = \alpha + \beta t + \gamma DI + \theta DS + \mu y_{t-1} + \sum_{i=1}^k c_i \Delta y_{t-i} + \varepsilon_t \quad \text{where } t = 1, \dots, T \quad (11)$$

DI is an indicator for the intercept shift and DS is the trend shift variable at the time of SB, where

$$DI_t = \begin{cases} 1 & \text{if } t > SB \\ 0 & \text{otherwise} \end{cases} \quad (12)$$

and

$$DS_t = \begin{cases} t - SB & \text{if } t > SB \\ 0 & \text{otherwise} \end{cases} \quad (13)$$

The coefficients of *DI* and *DS* reveal the dynamics of inequality in the short run and long run, respectively.¹¹ For example, a positive shift in the intercept occurs together with a negative shift in the trend coefficient. It suggests that inequality is rising in the short run, but mitigating in the long run.

The timing of *SB* is searched over the range of the sample $0.15T-0.85T$, where T is the sample size. Thus, this technique does not permit the structural break to take place in the beginning or at the end of the sample period. The test procedure first assigns a dummy variable for each of the possible break dates and then finds out the date that supports the alternative hypothesis the most. More specifically, the date with minimum t-statistics is chosen as the break date, i.e. $\text{argmin}_{SB} t_\mu(SB)$. The asymptotic distribution of the t-statistics used in this test does not follow the traditional *t*-distribution. The observation-specific finite sample critical value for the t-statistics is obtained from *ZA*, which is generated using a Monte Carlo simulation with the exact number of observations of above models. The null hypothesis of unit root is $\mu = 0$ and the alternative is stationary with one structural break. The lagged differenced terms are used to address the problem of serial correlation. Since the lag order is crucial for the size distortion, we follow Ng and Perron (1995) and use *t* tests to determine the lag order rather than using information criteria such as *AIC* and *BIC*.

However, a similar situation can happen for unit root tests with one structural break when the true DGP is trend stationary with two structural breaks. It is particularly interesting for our case since there are several abrupt events in China during the sample period. Thus, we proceed with the test allowing for two breaks proposed by Lumsdaine and Papell (1997). The model only permits shifts in the intercept (*LP1*):

$$\Delta y_t = \alpha + \beta t + \gamma DI1 + \omega DI2 + \mu y_{t-1} + \sum_{i=1}^k c_i \Delta y_{t-i} + \varepsilon_t \quad \text{where } t = 1, \dots, T \quad (14)$$

The model allows shifts in both intercept and slope (*LP2*):

$$\Delta y_t = \alpha + \beta t + \gamma DI1 + \theta DS1 + \omega DI2 + \psi DS2 + \mu y_{t-1} + \sum_{i=1}^k c_i \Delta y_{t-i} + \varepsilon_t \quad \text{where } t=1, \dots, T \quad (15)$$

DI1 and *DI2* are indicators for intercept shifts at time *SB1* and *SB2* with

$$DI1_t = \begin{cases} 1 & \text{if } t > SB1 \\ 0 & \text{otherwise} \end{cases} \\ DI2_t = \begin{cases} 1 & \text{if } t > SB2 \\ 0 & \text{otherwise} \end{cases} \quad (16)$$

¹¹ One of our referees suggests employing a smooth transition model such as Harvey and Mills (2004). However, in this paper, we focus on the timing of structural change in the economy. Moreover, we suspect the economy responded to policy rapidly because (1) the policies were usually unexpected and (2) the economy had resources reacting to the new situation, i.e., the labour was in abundance. Future research can use smooth transition models to examine the adjustment paths of different structural changes revealed in this paper.

DS1 and *DS2* are trend shift variables

$$\begin{aligned}
 DS1_t &= \begin{cases} t - SB1 & \text{if } t > SB1 \\ 0 & \text{otherwise} \end{cases} \\
 DS2_t &= \begin{cases} t - SB2 & \text{if } t > SB2 \\ 0 & \text{otherwise} \end{cases}
 \end{aligned} \tag{17}$$

The structural breaks are searched over a wider range from $0.1T$ to $0.9T$. Moreover, the procedure does not allow two breaks too close to each other by imposing the conditions $SB2 > SB1+1$ and $SB2 > SB1+2$ for model *LP1* and *LP2*, respectively. The lagged differenced terms are used to address the problem of serial correlation. The null hypothesis is $\mu = 0$ and the alternative is stationary with two structural breaks.

5 Empirical results

In the course of post-1949 economic development, China successively experienced many tremendous changes in its economic structure, which affected inequality remarkably. By matching the historical facts and endogenous breaks in the data, we look for the root of the structural breaks and examine the relationship between the regional inequality and the government policies. In this section, we explore the implications from the tests' results.

5.1 Empirical results: *ZA*

The results of the models *ZA1* and *ZA2* are summarised in Table 2 and Table 3, respectively. In both models, the null hypothesis of a unit root is rejected for each inequality measure. The rejections are much stronger for model *ZA2* where changes in both the intercept and slope are allowed. If connecting these endogenous break points to the historical events, the break dates obtained from the tests reveal very meaningful information.

<Table 2 and Table 3 about here>

For model *ZA1*, the upward shift in the level of the four inequality measures is found during 1990–1994, which is the period between the end of the second stage of the reform and the beginning of the third stage. Compared with the other measures, the inland and coastal inequality experienced the largest jump with $\gamma = 0.4217$ in the earliest date 1990. Because of the preferential economic policies and the geographical advantages, the coastal region attracted most FDI at the very early stage. In particular, Shanghai was granted the right to develop the Pudong New Area in 1990, which further deepened the reform in the coastal area, in turn, leaving the inland areas more economically behind.

For model *ZA2*, the breaks happened in the first stage of the reform during 1979–1984 except for GE, in 1990. It reveals another version of the reality during the transformation of the Chinese economy. As mentioned before, this period was the transformation period - from central planning economy to a gradually market-oriented economy. The household responsibility system was first introduced in the rural area and, later, the artificially controlled low price on agricultural products was relaxed step by step. The income for rural households increased due to better incentives and high market prices for agricultural products. It helps to close the gap between rural

and urban incomes, as shown in our result that a downward shift in constant ($\gamma = -0.2519$). Quickly following the rural reform, the government deepened the reform to include the urban areas, which altered the trend in urban economic development. In turn, it contributes to the divergence in the growth rate between the rural and urban areas in the following years, and therefore, causes an upward trend in the rural-urban inequality as shown in Table 3, $\theta = 0.0192$.

The break in the Gini coefficient occurred at the end of the first stage of the reform in 1983 whereas the GE rose very rapidly after 1990. The consumption inequality is consistent with the sigma divergence in income shown in Yao and Zhang (2001) using data from 1978 to 1995. Using different measures, it confirms that the inequality in China is indeed widening since the first stage of the reform.

5.2 Empirical results: LP

Table 4 reports the results for the model *LPI*. In contrast to the model *ZAI*, the null hypotheses cannot be rejected for all the inequality measures. Ben-David, Lumsdaine and Papell (1997) point out that allowing for additional breaks does not necessarily produce more rejections of the null hypothesis of unit root because the absolute critical values increase if more breaks are included. We therefore view the non-rejections due to the t-statistic on μ do not raise enough to compensate for the loss in power caused by the additional break in intercept.

<Table 4 about here>

Our results in Table 5 show that it is important to allow breaks in both the intercept and slope of the trend function. In contrast to model *LPI*, the empirical results for model *LP2* show that every null hypothesis is rejected at 1% significance level even though the absolute critical values increase from the previous models. Moreover, none of the coefficients for the second breaks in the intercept ω is significant at 5% level, which suggests that only allowing two breaks in the intercept may mis-specify the DGP of Chinese inequality measures.

<Table 5 about here>

Allowing for two breaks in both the intercept and trend produces richer results on the structural changes.¹² For each inequality measure, the first break happened in the pre-reform period and the other in the post-reform period. The first break in the inland-coastal inequality occurred in 1966 in which the break in trend is positive and significant ($\theta = 0.0734$). On the other hand, the timing of first breaks for the Gini, GE and rural-urban inequality coincided with the end of the Cultural Revolution during 1975–1976. Comparing the coefficients for the first break in intercept γ and slope θ for these measures, it illustrates that the shifts in intercept are positive whereas the

¹² In response to the suggestion of one of our referees, we ran the *LP2* model for two sub-samples with 30 observations to examine any other possible breaks in the sample period. One sample includes the first 30 observations and the other includes the last 30 observations. Using the first sample, both breaks occur two years around the SB1 in the full sample analysis which does not indicate further breaks other than the SB1. On the other hand, using the latter sample, there is a break happening in the mid-1990s in addition to the one in the mid-1980s. Although the results are indicative, the observations from these tests are too few to draw a definite conclusion. We suggest future research can employ unit root tests allowing for multiple breaks to analyse the significance of this possible break in a unified framework.

shifts in slope are negative. As shown in Figure 1, the fitted line shows that at the point of the first break, there is a jump in inequality—the positive γ s; after this point until to the next break point, the inequality measures are declining over time. The initial jump was caused by renewing the old policies before the Cultural Revolution, which focused on heavy industry in urban area and assigned low prices to agricultural products. However, as advocated in Lin et al. (2003), the policy cannot persist since it contradicts the comparative advantages of China which favours a labour-intensive agrarian economy. Therefore, in the long run, the income gap tended to be closed until the next policy struck and changed the fundamental economic development process.

The second break date for the inland-coastal inequality happened in 1981 and that for rural-urban inequality, Gini and GE occurred in 1986. The timing of structural breaks matches with the reform strategy. First, the SEZs were set up in 1981 by the open-door policy, which led to a prompt response of the inland-coastal inequality. Jones et al. (2003) show that preferential policies awarded to SEZs enhances growth by 5.5% per year. Second, in the second stage of the reform, the government shifted the focus from the rural to the urban. The positive coefficient of the rural-urban inequality, $\psi=0.1153$, signals the upward trend in the widening gap between the rural and the urban areas after the break point. Third, the increasing inequalities between inland and coastal areas and between rural and urban areas, together, led to a rising overall inequality of Gini/GE, which caused their second break in 1986.

Our results illustrate three aspects of inequality dynamics. First, the responses of inequalities to policy are prompt in China. The timing of structural breaks in the data is close to episodic events, which suggests minimal policy lag. Owing to the development strategy, the inland-coastal inequality usually reacts earlier than the other inequality measures. Second, our empirical results of consumption inequality are consistent with the literature of income convergence; with a mild decline in regional inequality found in Tsui (1991) using data from 1952 to 1985. Raiser (1998) shows that income inequality is reduced over the period 1978–1992, but the rate of convergence declines since 1985. By applying unit root tests with endogenous structural breaks to an extended sample period, we show that the inequality indeed increased after 1986. Third, our results show that the Chinese inequality evolution follows the pattern of income inequality in the U.S. shown in Coughlin and Mandelbaum (1988). A common driving force behind the similar pattern of these two countries is due to the reallocation of firms to developed areas, i.e. to the northeastern part of the US and to the coastal areas in China.¹³

6 Conclusions

The transition from a planned economy to a market economy in China resulted in remarkable economic growth. This article investigates one dimension of the economic development: rising regional inequality. We find that the regional inequality measures are trend stationary with structural breaks. The presence and timing of breaks in our data are consistent with the historical events of economic development in China. Matching the historical facts and the endogenous breaks in the data, we find that the Chinese government policies are very influential and have long-lasting effects on

¹³ The agglomeration effect might be less in China compared with that in the U.S. since Chinese policies on migration restricts people moving from backward region to an economic prosperous region freely (Au and Henderson 2006).

inequality.¹⁴ If no measures are taken to deal with this unfavourable trend, the regional inequality will continue to deteriorate in the future.

The ongoing plan of Western Region Development, which was launched in 1999 by the Chinese government, aims to speed up the development in the inland to catch up with its coastal counterpart. If this policy causes a fundamental change in the economic growth pattern in the inland area, it will lead to another structural break in the inequalities. However, in its first four years, Ho and Li (2008) do not find any obvious evidence for the effectiveness of this plan in helping the poor cities in the west out of their current status and therefore, reversing the trend of per capita output inequality between inland and coastal regions. Future research on the inequality may focus on examining and quantifying the effects of policy experiments on the economic structure.

¹⁴ Our findings are consistent with those in Ying (2007), which shows that institutional effects play an important role in the Chinese regional development.

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Table 1. Unit root tests without structural breaks

Variables	ADF	PP	DF-GLS
Gini	-2.292 (2)	-2.168 (3)	-1.667 (1)
GE	-1.698 (2)	-1.714 (3)	-1.369 (1)
Rural-Urban	-2.001 (2)	-2.061 (3)	-1.872 (1)
Inland-Coastal	-0.925 (6)	-2.577 (3)	-0.706 (0)

Notes: All variables are transformed to natural logs. The listed values are the t statistic of the test.

Null hypothesis is that $\mu = 0$ (there is a unit root). The numbers in parentheses for column 2 and 4 are the lag order in the ADF and DF-GLS tests. The lag orders for ADF are selected based on the Akaike's Information Criterion (AIC) and those for DF-GLS selected according to the Modified AIC (Ng and Perron 1995). The truncation lags are in the parentheses for PP test (column 3), which are Newey-West lags here, specified as $\text{int}\{4*(N/100)^{(2/9)}\}$ lags. Critical values for DF-GLS tests are from Elliot et al. (1996). All tests include breaks in both trend and constant in the model. These three tests are finished in STATA 8.2.

Table 2. The results for the model ZA1

Variables	SB	t_{μ}	γ	k
Gini	1994	-4.1892**	0.0891(3.1535)*	6
GE	1991	-3.5795***	0.1171(2.7396)*	3
Rural-Urban	1994	-3.7623**	0.1177(2.4277)*	3
Inland-Coastal	1990	-4.3323*	0.4217(3.2876)*	3

Notes: All variables are transformed to natural logs. t_{μ} is the t statistic of the test. Null hypothesis is that $\mu = 0$ (there is a unit root). For estimated time of break relative to total sample size, 4/5 here, the critical values are -4.33 at the 1% level, -3.75 at the 5% level and -3.46 at the 10% level (Zivot and Andrews 1992, p.256, table 2). The numbers in the parentheses are the t-statistics for the estimated coefficient. The lag length is chosen according to Ng and Perron (1995). *, ** and *** indicate significance at the 1%, 5% and 10% levels, respectively.

Table 3. The results for model ZA2

Variables	SB	t_{μ}	γ	θ	K
Gini	1983	-6.0444*	-0.1044(-3.5791)*	0.0165(5.6233)*	6
GE	1990	-5.4569*	-0.0136(-0.2671)	0.0440(4.3887)*	6
Rural-Urban	1981	-5.5101*	-0.2519(-4.0182)*	0.0192(4.5292)*	3
Inland-Coastal	1982	-5.7108*	-0.3289(-3.0361)*	0.0701(4.7761)*	3

Notes: All variables are transformed to natural logs. t_{μ} is the t statistic of the test. Null hypothesis is that $\mu = 0$ (there is a unit root). For estimated time of break relative to total sample size, $7/10$, the critical values are -4.75 at the 1% level, -4.18 at the 5% level and -3.86 at the 10% level (Zivot and Andrews 1992, p.257, table 4). The numbers in the parentheses are the t -statistics for the estimated coefficient. The lag length is chosen according to Ng and Perron (1995). * indicates significance at the 1% level.

Table 4. The results for model LP1

Variables	SB1	SB2	t_{μ}	γ	ω	k
Gini	1962	1979	-5.3209	-0.2025(-4.8803)*	-0.1511(-4.8801)*	6
GE	1991	1995	-4.3219	0.1125(2.2895)**	0.1499(2.7142)**	5
Rural-Urban	1981	1992	-5.0718	-0.1787(-3.0014)*	0.1768(3.7251)*	3
Inland-Coastal	1987	1991	-4.8602	0.2782(2.6243)**	0.3346(2.5231)**	3

Notes: All variables are transformed to natural logs. t_{μ} is the t statistic of the test. Null hypothesis is that $\mu = 0$ (there is a unit root). The critical values are -6.74 at 1% level, -6.16 at 5% level and -5.89 at 10% level from Ben-David, Lumsdaine and Papell (1997). Critical values were computed for model only with shift in intercept using 125 observations and 5,000 replications. They use the method of endogenously determining the appropriate lag length in computing critical values, with a maximum of 8 lags. The data were generated under the null hypothesis, $y_t - y_{t-l} = \varepsilon_t$. The values are slightly different from those reported in Lumsdaine and Papell (1997) since the number of replications increased from 500 to 5,000. The numbers in the parentheses are the t -statistics for the estimated coefficient. The lag length is chosen according to Ng and Perron (1995). * and ** indicate significance at the 1%, and 5% levels, respectively.

Table 5. The results for model LP2

Variables	SB1	SB2	t_{μ}	γ	ω	θ	ψ	k
Gini	1974	1986	-9.8971*	0.1124(5.1242)*	-0.0262(-1.2531)	-0.0121(-4.7796)*	0.0407(8.7168)*	6
GE	1975	1986	-8.7677*	0.2167(4.7541)*	-0.0747(-1.6779)	-0.0305(-5.0300)*	0.0913(8.0053)*	6
Rural-Urban	1975	1986	-7.4130*	0.2960(4.5499)*	-0.1152(-1.7838)***	-0.0564(-5.8123)*	0.1153(6.8484)*	6
Inland-Coastal	1966	1981	-8.4610*	0.0997(0.8770)	-0.5389(-4.6986)	0.0734(3.4432)*	0.0649(5.2113)*	2

Notes: All variables are transformed to natural logs. t_{μ} is the t statistic of the test. Null hypothesis is that $\mu = 0$ (there is a unit root). The critical values are -7.19 at 1% level, -6.75 at 5% level and -6.48 at 10% level from Ben-David, Lumsdaine and Papell (1997). Critical values were computed for models with shift in both intercept and trend using 125 observations and 5000 replications. They use the method of endogenously determining the appropriate lag length in computing critical values, with a maximum of 8 lags. The data were generated under the null hypothesis, $y_t - y_{t-1} = \varepsilon_t$. The values are slightly different from those reported in Lumsdaine and Papell (1997) since the number of replication increased from 500 to 5,000. The numbers in the parentheses are the t -statistics for the estimated coefficient. The lag length is chosen according to Ng and Perron (1995). * and *** indicate significance at the 1% and 10% levels, respectively.

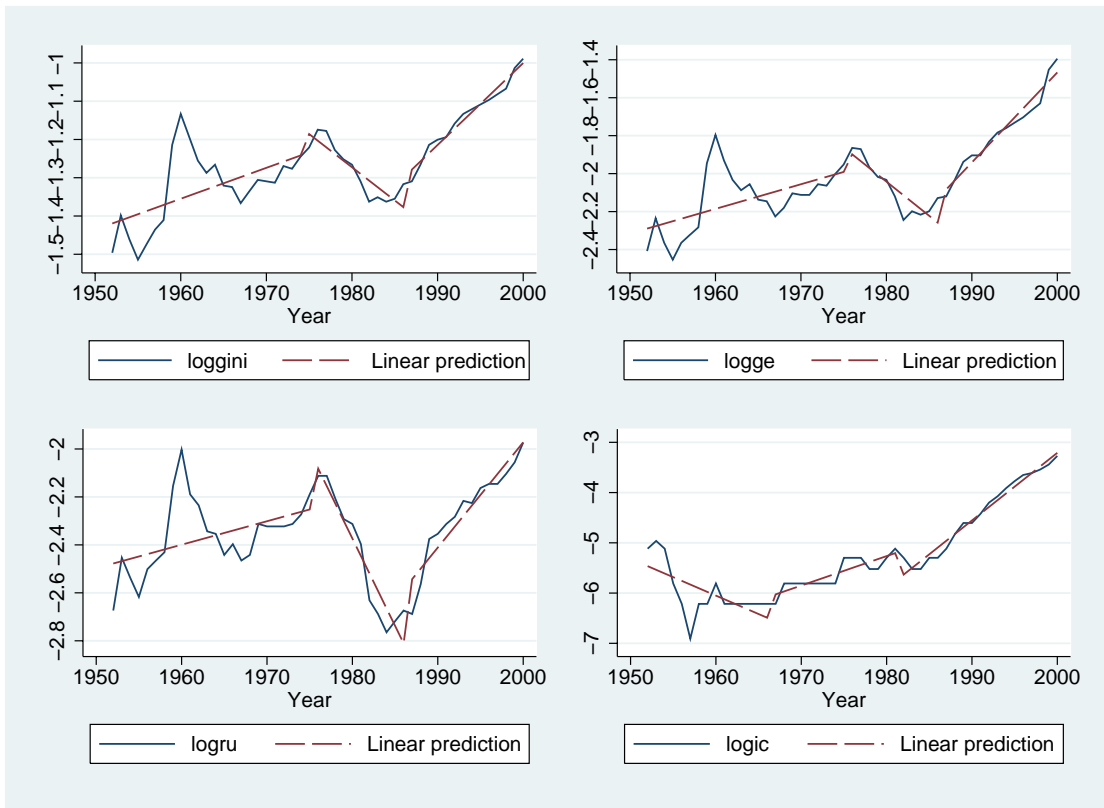


Fig. 1. The log of Gini, GE, rural-urban & inland-coastal inequality measures and their fitted values, 1952–2000

Appendix

Table A1. Inequality measures in China, 1952-2000

Year	Gini (%)	GE (%)	Rural-Urban (%)	Inland-Coastal (%)
1952	22.4	9	6.9	0.6
1953	24.7	10.7	8.6	0.7
1954	23.2	9.4	7.9	0.6
1955	22	8.6	7.3	0.3
1956	22.9	9.4	8.2	0.2
1957	23.8	9.8	8.5	0.1
1958	24.4	10.2	8.8	0.2
1959	29.7	14.3	11.6	0.2
1960	32.2	16.6	13.5	0.3
1961	30.3	14.5	11.2	0.2
1962	28.5	13.1	10.7	0.2
1963	27.6	12.4	9.6	0.2
1964	28.2	12.8	9.5	0.2
1965	26.7	11.8	8.7	0.2
1966	26.6	11.7	9.1	0.2
1967	25.5	10.8	8.5	0.2
1968	26.3	11.3	8.7	0.3
1969	27.1	12.2	9.9	0.3
1970	27	12.1	9.8	0.3
1971	26.9	12.1	9.8	0.3
1972	28.1	12.8	9.8	0.3
1973	27.9	12.7	9.9	0.3
1974	28.8	13.5	10.3	0.3
1975	29.5	14.2	11.2	0.5
1976	30.9	15.5	12.1	0.5
1977	30.8	15.4	12.1	0.5
1978	29.3	14	11	0.4
1979	28.6	13.3	10.1	0.4
1980	28.2	13.1	9.9	0.5
1981	27	12	9.1	0.6
1982	25.6	10.6	7.2	0.5
1983	25.9	11.1	6.8	0.4
1984	25.6	10.9	6.3	0.4
1985	25.8	11.1	6.6	0.5
1986	26.8	11.9	6.9	0.5
1987	27	12	6.8	0.6
1988	28.2	13.1	7.7	0.8
1989	29.7	14.4	9.3	1
1990	30.1	14.9	9.5	1
1991	30.3	14.9	9.9	1.2
1992	31.4	16	10.2	1.5
1993	32.2	16.8	10.9	1.7
1994	32.6	17.2	10.8	2
1995	33	17.7	11.5	2.3
1996	33.4	18.2	11.7	2.6
1997	33.9	18.9	11.7	2.7
1998	34.4	19.6	12.2	2.9
1999	36.3	23.4	12.8	3.2
2000	37.2	24.8	13.9	3.8

Source: Kanbur and Zhang (2005)