Understanding regional growth dynamics in Japan: panel cointegration approach utilizing the PANIC method

Masahiko Shibamoto[a,*], Yoshiro Tsutsui[b] Chisako Yamane[c]

[a] *Research Institute for Economics and Business Administration, Kobe University*

[b] *Graduate School of Economics, Osaka University*

[c] *Faculty of Economics, Niigata Sangyo University*

January 20, 2011

Abstract

This study aims at understanding regional growth dynamics in Japan using nonstationary panel data. Since the panel unit root test does not adequately produce a detailed picture of the development of Japanese prefectures, we follow a panel cointegration approach using the PANIC method. We find that there is one common source of growth to which prefectures attach different long-run weights and that the per capita real income of follower-prefectures will catch up to that of leader-prefectures. Using the concept of relative convergence, we find that although the poor stay poor, the relative income gap will narrow substantially in the future.

JEL Classification: O40, C22, C23

Keywords: β-convergence, Common trends, Panel unit root test, PANIC method, Relative convergence, Japan

[*]Correspondence to: Masahiko Shibamoto, Research Institute for Economics and Business Administration, Kobe University, 2–1, Rokkodai, Nada, Kobe, 657–8501, Japan. E–mail: shibamoto@rieb.kobe-u.ac.jp.

1 Introduction

In this study we develop a method for analyzing the convergence problem and apply it to regional economies in Japan. This paper focuses on β-convergence, a concept that is extensively developed and widely used.[1] β-convergence means that poor economies tend to grow faster than rich economies, or in terms of growth theory, the presence of long-run balanced growth paths are parallel across economies.[2]

Barro regression, a cross-sectional regression of the long-term per capita income growth rate on initial per capita income, is the first method to investigate β-convergence. It requires data only at the two remote points of time needed for testing the convergence hypothesis derived from the Solow model, and seeks empirical evidence of a negative correlation between the initial per capita income and its growth rate (Barro, 1991).

Many cross-sectional regression studies, including Barro and Sala-i-Martin (1992a,b), argue that follower-economies catch up to leader-economies at the annual rate of about 2% in the context of absolute or conditional convergence. Barro regression has the advantage of being parsimonious, but this advantage is at the same time a problem: it disposes of numerous data between the two remote points of time and tells us nothing about the dynamic process of growth. Another problem is that a negative correlation is a necessary condition for the convergence, so that convergence is not warranted even if a negative correlation is found in Barro regressions.

Bernard and Durlauf's (1995) Definition 2.1 and Evans and Karras (1996) provide a test for absolute and conditional convergence hypotheses using a unit root test on panel data (hereinafter the Evans and Karras test). Their idea is that two economies will converge if the difference in per capita income between the two economies is stationary. This condition

[1] For an overview of literature on convergence, including empirical papers using the panel data approach, see Durlauf et al. (2005).

[2] In the neoclassic growth theory, if each country has access to the same aggregate production functions (decreasing returns to production factors), the steady-state is independent of an economy's initial capital, labor stocks, and initial income. Long-run differences in output reflect differences in the determinants of accumulation, not differences in the technology used. Therefore, poor economies grow faster than rich ones, and the poor will eventually catch up with the rich. This type of convergence is called "absolute convergence". Even if one relaxes the assumption that countries have access to the same production functions, convergence in growth rates can still occur so long as each country's production function is concave in capital per efficiency unit of labor and each country experiences the same rate of labor-augmenting technological change. In such a case, although the gap between poor and rich economies shrinks over time, it does not completely vanish in the steady states. This type of convergence is called "conditional convergence".

neatly matches the definition of convergence. In addition, they argue that the panel unit root test has greater statistical merit for increasing the efficiency of estimations than the cross-sectional regressions.

However, when we examine whether there are common components affecting per capita incomes that differ in magnitude across economies, the above convergence definition becomes ambiguous. The possibility exists that the economies have access to heterogeneous technology and thus nonparallel long-run balanced growth paths may emerge. We can generalize the panel unit root test to allow that per capita incomes of economies cointegrate with common components with different long-run weights. In other words, the panel unit root approach is a special case of cointegration. This approach was proposed by Bernard and Durlauf (1995) and enables us to examine the short-run dynamic behavior of deviations from long-run equilibrium paths.

In this paper we follow the cointegration approach since it is the most comprehensive method today. Then, we face the problem of which statistical tool to use. The well known Johansen (1995) test, which is employed by Bernard and Durlauf (1995), seems a good candidate since it reports the cointegration rank and cointegration vectors as well as information on short-run dynamics from error correction terms. Suppose however, that we want to analyze long-run equilibrium growth among 50 countries. Even in the lucky event that annual data for 50 years are available, the degree of freedom is not adequate to estimate a Johansen model.[3] Thus, we need to develop a statistical tool that allows us to follow the cointegration approach. This paper achieves this goal by adopting the Panel Analysis of Nonstationarity in Idiosyncratic and Common Components (PANIC) method developed by Bai and Ng (2004).[4] This tool determines common factors of per capita incomes of individual economies before estimating the cointegration relation between per capita income and common factors.[5]

We apply the method to panel data from 46 Japanese prefectures from 1955 to 1999. Our analysis is divided into three stages. First, we explore the long-run equilibrium

[3] Bernard and Durlauf (1995) carried out a Johansen test with annual data from 1900 to 1987 for 15 countries, assuming that the lag-length is two.

[4] Pesaran (2007) propose a pairwise approach to testing for cointegration for all possible pairs of output gaps across economies. This method may be an alternative to the method adopted in this paper. His approach is applicable when the number of economies is large relative to the time dimension of the panel.

[5] Westerlund et al. (2010) Chinese economies using a similar method.

growth across the prefectures. Literature using panel data has not reached consensus on income convergence across Japanese prefectures.[6] While panel unit root tests have generally accepted the no convergence null hypothesis (e.g., Kawagoe, 1999), tests based on the dynamic panel regression approach, that is an extension of the Barro regression to the panel case, show that Japanese prefectures are converging at a rate faster than 2% annually (e.g., Shioji, 2001).[7] This paper provides a clear picture of the long-run per capita income of Japanese prefectures. Second, we analyze short-run dynamics by looking at the deviation from equilibrium paths. In analyses of both long-run equilibrium and short-run dynamics, we examine cross-prefectural properties to highlight the reality. Finally, we use the concept of relative convergence proposed by Phillips and Sul (2007) to analyze steady states in the future.

The remainder of the paper is structured as follows. Section 2 explains how to construct real per capita income data of the prefectures and conduct preliminary analyses. In section 3, the Evans and Karras type convergence hypothesis is tested using panel unit root tests. In section 4, we develop a framework for the cointegration approach adopting the PANIC method and empirically examine long-run and short-run dynamics of prefectural per capita income. Section 5 investigates steady states to determine whether poor prefectures will remain poor in the future. Concluding remarks are given in the final section.

2 Data and preliminary analyses

The only data necessary for our analysis are panel data of prefectural real per capita income. Our data are from 1955 to 1999 for 46 prefectures.[8] We exclude Okinawa Prefecture because it was occupied by the USA until 1972. Prefectural income data were taken from the *Annual Report on Prefectural Accounts* published by the Economic and Social Research Institute (ESRI) Cabinet Office. To obtain real prefectural per capita income,

[6]Current studies on convergence for countries around the world have not reached consensus either. See Islam (1995, 1998), Lee et al. (1997, 1998), and Evans and Karras (1996) for supporting evidence and Bernard and Durlauf (1995) and Quah (1996) for evidence against..

[7]Some empirical studies, such as Islam (1995, 1998), Lee et al. (1997, 1998), Shioji (2001), basically regress panel data regarding per capita income over its lagged value. We call this type of analysis the dynamic panel regression approach. Shibamoto and Tsutsui (2011) argue that this approach is problamatic.

[8]Prefectural income data prior to 1999 are based on the System of National Accounts of 1968 (68SNA). Althouh 93SNA data are available after 1999, we should not use them, because these two series are based on different definitions of the income and outlay account and its calculation methods. Therefore, to avoid bias, we analyzed the period that the 68SNA data cover.

total prefectural income was divided by the population of the prefecture as published in the *Population Census* and further divided by the prefectural price index.

Panel data for prefectural price indices are calculated by multiplying the *Regional Difference Index of Consumer Prices*, which reflects cross-sectional differences, by the *General Index of Tokyo*, which reflects time-series differences in the Tokyo area. [9] The pre-World War II period average of 1934–1936 is normalized at 1.

In 1955, the log per capita real income averaged 5.55 with a standard deviation (SD) of 0.17. By 1999, the average rises to 7.44, while the SD decreases to 0.10. The decrease in SD suggests that the income gap among prefectures narrows over these 45 years in the sense of ?-convergence. This supposition is also confirmed by a scatter plot of the log per capita income in the initial (1955) and final (1999) years of the study (Figure 1). The log per capita income is more widely dispersed in 1955 than in 1999. The plot also indicates a stronger relationship between per capita income in 1955 and 1999, implying that poor prefectures remained poor and rich prefectures remained rich.

Finally, the scatter plot in Figure 2 shows the log per capita real income in 1955 and annual growth rates for 1955–1999. The annual growth rates range 3.5–4.7%. The fitted line on the scatter plot (i.e., the Barro regression) shows a strong negative correlation.

These three preliminary analyses indicate that the income gap among Japanese prefectures has been decreasing since World War II. Since the Japanese economy was completely destroyed during the war and rebuilt from scratch after the war, it serves as a good exercise for the convergence problem. However, to learn how per capita income moves across prefectures, and to elucidate whether the prefectures will converge to the same per capita income level in the future, we need to rely on more sophisticated methods.

3 Regional growth convergence

Following Evans and Karras (1996), this section describes the panel unit root as a vehicle for testing β-convergence, then applies it to the Japanese prefectures.

[9] GDP deflator by prefecture lacks observations for some prefectures, so that we did not use them.

3.1 Time–series convergence tests using panel data

Consider a collection of economies $1, 2, \cdots, N$ that have eventual access to the same body of technical knowledge, implying that the balanced growth paths of the N economies are parallel. Evans and Karras (1996) show that economies i and j converge if

$$\lim_{l \to \infty} (y_{i,t+l} - y_{j,t+l}) = \mu_{i,j}, \text{ for every pair of economies } i \text{ and } j, \tag{1}$$

and equivalently,

$$\lim_{l \to \infty} (y_{i,t+l} - \bar{y}_{t+l}) = \mu_i, \text{ for all } i = 1, \cdots, N, \tag{2}$$

where y_{it} is the logarithm of prefectural per capita real income, $\bar{y}_t \equiv (\sum_{i=1}^{N} y_{it})/N$, and μ_i is a permanent cross-economy difference.[10] The condition that countries converge along parallel equilibrium growth paths is represented by economy-specific intercepts μ_i. If the absolute convergence hypothesis holds, then $\mu_i = 0$ for all i.

Evans and Karras (1996) show that equation (2) can be tested with panel data by estimating

$$y_{it} - \bar{y}_t = \mu_i + \rho_i (y_{i,t-1} - \bar{y}_{t-1}) + u_{it}, \text{ for all } i = 1, 2, \cdots, N, \tag{3}$$

where u_{it} is the error term with a zero mean for economy i. Whether or not convergence occurs is evaluated by testing if the value of the autoregressive parameter ρ_i is unity.

To test the stationarity of the income differentials, $y_{i,} - \bar{y}_t$, we employ pooled tests based on Fisher痴 statistic, as defined in Choi Choi (2001). Choi痴 Choi (2001) test combines p-values from a cointegration test applied to each economy under the null hypothesis that all cross section units have a unit root, against the alternative that some of all panel units are stationary. Choi痴 Choi (2001) test statistic, which we call PN statistic, is as follows:

$$P_N = \frac{1}{2\sqrt{N}} \sum_{i=1}^{N} (-2 \ln p_i - 2) \overset{d}{\Rightarrow} N(0,1) \text{ as } T \to \infty, N \to \infty \tag{4}$$

where p_i is the p-value associated with unit root/cointegration test statistic for economy

[10] Bernard and Durlauf (1995) call this Definition 2.1.

i.

We use two types of unit root tests with a constant: the ADF test where the null hypothesis is a unit root (Dickey and Fuller, 1979) and the KPSS test where the null hypothesis is stationarity (Kwiatkowski et al., 1992).[11] Because the power of the tests of the unit root null is low in small samples, testing the stationarity null is indispensable.

3.2 Empirical results of the Evans and Karras convergence test

The second column of Table 1 presents a Choi (2001) statistic P_N based on the ADF type unit root test for the individual prefectures.[12] This allows us to test the null hypothesis that all prefectures have a unit root versus the alternative that some fraction is stationary. The result indicates that we can reject the null hypothesis that all prefectures have a unit root at the 5% significance level.

The third column of Table 1 presents a Choi (2001) statistic P_N based on a KPSS type unit root test for the individual prefectures.[13] This allows us to test the null hypothesis that all prefectures are stationary versus the alternative that some of them have a unit root. The result indicates that we can reject the null hypothesis that all prefectures are stationary at the 5% significance level.

The results of the two types of panel unit root tests indicate that some prefectures are stationary and others have a unit root, which may imply that some prefectures converge and others do not. This supposition can be examined by inspecting the results of the unit root test for each prefecture before aggregating them to Choi's P_N, which are presented in Table 2. The table reveals that prefectures such as Akita, Ishikawa, Nagano, Toyama, Shimane, and Kagawa are stationary, while around half of the prefectures including Tokyo, Kanagawa, Osaka, and Kyoto have a unit root.[14] Of course, it is too hasty to conclude that the former group converges and the latter group does not, because the measuring deviations from the average prefectural income $\bar{y}_t$ are questionable for this interpretation.

[11] The KPSS test examines whether the variance of the stochastic trend component of the series is zero.

[12] The number of lags of the lagged difference terms of income differentials was set at four.

[13] The number of truncation lags in the KPSS test was set at 12.

[14] This result is similar to the finding of Pesaran (2007) who apply the pair-wise approach to all possible pairs of output series in the Penn World Tables over 1950–2000, and obtain the estimates of the proportion of output pairs for which convergence hypothesis is not rejected. He found that there are some pairs for which the differences in output are stationary with a constant mean, although there is less evidence of output convergence at a global level.

However, the results do imply that the autoregressive coefficient ρ_i is not identical across prefectures. Indeed, we tested whether the autoregressive coefficient ρ_i is identical for all prefectures and the null hypothesis of identical ρ is rejected at the 1% significance level, as shown in the fifth column of Table 1.

In conclusion, if we follow the panel unit root test approach, empirical results suggest that Japanese prefectures do not converge along parallel equilibrium growth paths.

4 Regional growth dynamics: common and prefecture-specific components

4.1 Nonparallel balanced growth paths

In the previous section, the results of panel unit root tests do not support the convergence hypothesis that all the prefectures have the same steady state growth rate. However, this does not necessarily mean that they have no common source of growth. The panel unit root approach focuses on a special case of long-run equilibrium relationship (i.e., convergence). Specifically, it requires the autoregressive coefficient of equation (3) to be homogeneous, implying that all prefectures have parallel balanced growth paths, that is, their convergence speed is identical. We have shown that Japanese prefectures do not have parallel balanced growth paths; however, there is some possibility of heterogeneous (or nonparallel) balanced growth paths.[15]

This section investigates a model allowing for balanced growth paths to differ across prefectures. In practice, we decompose the per capita income into common and prefecture specific components, and examine whether each per capita income cointegrates with the common components but with different long-run weights. Such an idea is not new, and Bernard and Durlauf's (1995) Definition 2.2 (hereafter BD2) of *common trends in output* indeed embodies this idea. Thus, we follow BD2 to elucidate if there are common trends among Japanese prefectures. However, our method is new in its use of the econometric framework proposed by Bai and Ng (2004). Use of their framework makes the cointegration approach practically possible and enables us to investigate both the long-run equilibrium

[15]Lee et al. (1997) and Pesaran (2007)) have rationalized this possibility in the context of a neoclassical growth model.

path and short-run deviations from it, for each prefecture.

4.2 Econometric framework based on the panel cointegration approach

Similar to BD2, we define that the economies $1, 2, \cdots, N$ contain common trends, if

$$\lim_{l \to \infty} (y_{it+l} - \lambda_i' F_{t+l}) = \mu_i, \tag{5}$$

where F_t is the r common trends followed by the economies, and λ_i is a parameter vector that represents different weights to the common trends.

Comparing equation (5) with equation (2) in Section 3, we see that the convergence hypothesis defined by Evans and Karras (1996) and Bernard and Durlauf's (1995) Definition 2.1 is a special case in which $\lambda_i = 1$ for all i and the mean value for per capita income $\bar{y}_t$ is a variable of common trends F_t. In contrast, by adopting equation (5) we allow for the possibility that heterogeneous balanced growth paths exist, and it becomes possible to examine whether there are common trends for all economies.

Equation (5) has a natural testable counterpart in unit root/cointegration literature. If $e_{it} \equiv y_{it} - \mu_i - \lambda_i' F_t$ is a mean zero stationary process, then equation (5) will be satisfied. In addition, equation (5) allows each economy y_{it} to respond to the common trends with a different weight λ_i which needs to be estimated. To achieve these estimations, we use Bai and Ng (2004)'s PANIC method, which first determines common components of income and then tests if the common components are nonstationary.

Specifically, Bai and Ng (2004) consider the factor model:

$$y_{it} = \mu_i + \lambda_i' F_t + e_{it}, \tag{6}$$

where e_{it} for $i = 1, \cdots, N$ is the idiosyncratic component with a zero mean and is orthogonal to F_t and to each other. Taking the first difference of equation (6) yields:

$$\Delta y_{it} = \lambda_i' f_t + \Delta e_{it}, \tag{7}$$

where $f_t \equiv \Delta F_t$. Applying the principal-components analysis to Δy_{it}, we obtain estimates of r factors $\hat{f}_t$. Then, calculating back $\hat{F}_t \equiv \sum_{s=2}^{t} \hat{f}_s$ for $t = 2, \cdots, T$ and conducting a least-squares fit of equation (6), we obtain the estimators of μ_i and λ_i and the residuals $\hat{e}_{it}$ for each $i = 1, \cdots, N$.

To test the stationarity of the common component $\hat{F}_t$, we use two types of unit root tests: the ADF test where the null hypothesis is a unit root, and the KPSS test where the null hypothesis is stationarity. In both tests, constant and time trends are included.

To test the stationarity of an estimated idiosyncratic component $\hat{e}_{it}$, Bai and Ng (2004) propose using pooled tests based on Choi (2001)'s statistic P_N. We apply two types of cointegration tests of equation (6): the Phillips and Ouliaris (1990) test where the null hypothesis is no cointegration and the Shin (1994) test where the null hypothesis is cointegration.

4.3 Empirical results using our framework: the long-run equilibrium path

This subsection provides empirical evidence obtained by the PANIC method. We first test whether all prefectures have long-run common trends and then investigate time series and cross-sectional properties for the long-run equilibrium growth paths and their short-run dynamics.

First, we report the Bai and Ng (2002)'s information criteria $IC(r)$ to determine the number of common factors r in equation (7).[16] The criteria is minimized at $r = 1$. The estimated factor F_t is depicted in Figure 3, which changes at the same rate as the mean value of the per capita real income across prefectures $\bar{y}_t$. Indeed, the two variables are highly correlated (the correlation coefficient is 0.99).

Next, we investigate the time series properties of the common component F_t and the prefecture-specific component $\hat{e}_{it}$ for all prefectures. The upper panel of Table 3 presents the unit root test results of the common component F_t. We find that the ADF test accepts the null hypothesis that the common component F_t has a unit root (the left

[16]To determine the number of common factors r in equation (7), Bai and Ng (2002) derive information criteria. We adopted one criterion of their six, that which is most robust when there is cross-correlation among the idiosyncratic components: $IC(r) \equiv \log(V(r,f)) + r(N+T)/(NT)\log(NT)$. The information criteria reflect the trade-off between goodness-of-fit and overfitting. The first term on the right shows the goodness-of-fit which is given by the residual sum of squares, $V(r,f) \equiv 1/(NT)\sum_{i=1}^{N}\sum_{t=1}^{T}(\Delta y_{it} - \lambda_i f_t)^2$, which depends on estimates of the factors and the number of factors. If the number of factors r is increased, variance of the factors f_t also increases, while the sum of squared residuals decreases. The penalty of overfitting, which is the second term on the right, is an increasing function of the cross-section size N and time series length T. The optimal number of factors minimizes $IC(r)$.

column) and the KPSS test rejects the null hypothesis that the factor is stationary at the 5% significance level (the right column), unequivocally indicating that the common factor F_t is a nonstationary process.

The lower panel of Table 3 presents the results on idiosyncratic components. We apply Choi (2001)'s statistic P_N based on the Phillips and Ouliaris (1990) cointegration test to each prefecture. The null hypothesis that all the estimated prefecture-specific components $\hat{e}_{it}$ are nonstationary is rejected at the 5% significance level (the left column). Further, the null hypothesis that all the estimated prefecture-specific components $\hat{e}_{it}$ are stationary was accepted by Choi (2001)'s statistic P_N based on the Shin (1994) cointegration test applied to each prefecture (the right column).[17]

In conclusion, prefectural per capita real income in Japan has one long-run equilibrium relationship, where the nonstationary behavior is driven by the common univariate time series F_t, which is almost identical to the average per capita income of Japan. However, each prefecture attaches a different weight λ_i to it, implying that the prefectures follow different long-run balanced growth paths.

4.4 Cross-prefectural properties of the long-run equilibrium growth path

The results of Table 3 demonstrate how the convergence hypothesis specified as equation (2) is not supported by the conventional panel unit root tests in Section 3. Given equation (6), the difference between two per capita real income series is denoted as $y_{it} - y_{jt} = \mu_i - \mu_j + (\lambda_i - \lambda_j)F_t + e_{it} - e_{jt}$. Given that F_t is univariate nonstationary if $\lambda_i \neq \lambda_j$, then y_{it} is not cointegrated with y_{jt}. Thus, the results from the conventional panel unit root tests in Section 3 that some pairs of prefectures are cointegrated with the prefecture mean $\bar{y}_t$, while others are not, really means that the factor loading coefficients coincide for some prefectures, but not for others. This argument is similar to that of Pesaran (2007).[18]

The long-run equilibrium is characterized by two parameters, μ_i and λ_i. While λ_i can be interpreted as the slope of an individual balanced growth path, reflecting cross prefecture differences in convergence speed, μ_i represents the prefecture-specific fixed ef-

[17] In conducting the Phillips and Ouliaris (1990) test, the lagged difference terms of $\hat{e}_{it}$ were added to capture serial correlations, in which the number of lags was set at four. Similarly, the number of truncation lags in the Shin (1994) test was set at 12.

[18] Pesaran (2007) argue that under $\lambda_i F_t \sim I(1)$ economies i and j converge in the sense of Evans and Karras (1996) and Bernard and Durlauf's (1995) Definition 2.1, if $\lambda_i = \lambda_j$.

fect. Scatter plots show a clear negative correlation between μ_i and λ_i (Figure 4). This result indicates that follower-prefectures, which started with a lower income represented by a smaller μ_i, have a relatively larger λ_i, implying that they are catching up to leader prefectures that started with a higher income. However, this does not mean that each prefecture converges to a common or parallel balanced growth path, because not all prefectures converge in the sense of equation (2), as is empirically shown in Section 3.

4.5 Cross-prefectural properties of the short-run dynamic behavior

We use five analyses to examine cross-prefectural properties of the short-run dynamic behavior. First, since it is difficult to draw clear conclusions from the examination of each pattern of 46 prefectures, we divide the 46 prefectures into two groups and analyze their characteristics. Specifically, we apply k-means cluster analysis to the prefecture-specific components $\hat{e}_{it}$ to divide the 46 prefectures into two groups so as to maximize the similarities of the constituents.[19] The colors in the map of Japan (Figure 5) differentiate between the two groups of prefectures as clusters 1 and 2. The Figure shows that cluster 1 includes metropolitan prefectures such as Tokyo, Aichi, and Osaka, while cluster 2 consists of rural prefectures such as those in the Tohoku and Kyushu regions.

Next, we examine how these two clusters differ with respect to the long-run growth rate. To this aim we depict the scatter plot of μ and λ (see Figure 4). μ_i reflects the initial income level and λ_i reflects the income growth rate in the long-run of prefecture i. As explained in the previous subsection, there is a negative relationship between λ and μ, implying that poor prefectures will grow faster and catch up with the current rich prefectures in the long-run. In Figure 4, cluster 1 is indicated by dark diamonds and cluster 2 by bright squares. In general, prefectures belonging to cluster 1 are located in the lower right of the figure and prefectures belonging to cluster 2 in the upper left, suggesting that cluster 1 represents the leader-prefectures and cluster 2 follower-prefectures. Interestingly, stratification of the clusters based on short-run behavior corresponds to stratification of clusters based on long-run behavior.

[19]Our measure of dissimilarity in a cluster is the squared Euclidean distance from the center of the cluster, $\sum_{c=1}^{2}\left\{\sum_{i\in c}\left[\sum_{t=1956}^{1999}(\hat{e}_{it}-\bar{\hat{e}}_{ct})^2\right]\right\}$, where $\bar{\hat{e}}_{ct}$ is the average for all $\hat{e}_{it}$ in cluster c. To select the best set of clusters from a number of clustering exercises with different initial centers, we applied the k-means algorithm 10,000 times to $\hat{e}_{it}$ with random initialization. We then searched the resulting 10,000 sets of clusters to determine the set that minimized the distance defined above.

Third, we show that poor prefectures catch up with rich prefectures at a higher rate during the recession periods. This is seen when we depict short-run deviations from the common trend, $\hat{e}_{it}$, for clusters 1 and 2 for the observation period in Figure 6 where recession periods, from their peaks to their troughs, are shown as solid grey background.[20]

The gap between the graphs, e_{it} of cluster 1 $-$ that of cluster 2, represents the short-run income gap between leader-prefectures and follower-prefectures.[21] Widening of the gap indicates that the leader-prefectures temporarily grew faster than the followers. A close inspection of Figure 6 reveals that the short-run gap narrows during recessions except between 1980 and 1983.

Fourth, we examine if the inequality of regional income reflects, at least partially, the short-run gap shown in Figure 6. Specifically, the inequality should become smaller when the follower-prefectures catch up rapidly (i.e., the short-run gap is smaller). To check if this is really the case, we calculate the Gini index of the logarithm of prefectural per capita income for each year from 1955 to 1999 (Figure 7). The correlation coefficient between the changes in the Gini index and those in the short-run gap is 0.784, implying that the Gini index is well synchronized with the increases and decreases of the short-run gap.

Finally, we examine whether our results of short-run dynamics are consistent with the change in speed of convergence.[22] Speed of convergence in the short-run is nothing but the rate of change in the gap between leader-prefectures (i.e., cluster 1) and the follower-prefectures (i.e., cluster 2). Thus, our hypothesis is that the speed of convergence is low as the gap between clusters 1 and 2 increases and high as it decreases.

In order to examine this hypothesis, we need to measure the convergence speed for subperiods. For the measurement, we split the sample into five-year periods and conduct cross-sectional regressions. We estimate the following regression model by the nonlinear least squares equation for each subperiod:

$$\frac{(y_{iT} - y_{i0})}{T} = \alpha - \frac{1 - \exp(-\beta T)}{T} y_{i0} + \epsilon_i, \tag{8}$$

where 0 denotes the initial year and T the final year for each subperiod. The coefficient β

[20] The data of peaks and troughs are taken from ESRI.

[21] The graphs are symmetrical by construction, since the total deviation from the average is zero.

[22] Applying Barro regression to Japanese regional economies, Barro and Sala-i-Martin (1992b) and Yamane and Tsutsui (2009) find that the equality of β coefficients between the short subperiods is rejected.

measures the speed of shrinkage of the gap between the initial and final years.

Table 4 presents the estimate of β for 1955–1999 and for the nine subperiods of five years each.[23] According to Table 4, the convergence speed was low for the periods 1955–1960, 1965–1970, and 1975–1990, and high for the periods 1960–1965, 1970–1975, and 1990–1999. This change in the speed of convergence in the short-run is in synchronization with the increases and decreases in the short-run gap between cluster 1 and cluster 2. Indeed, the scatter plots (Figure 8) of the estimates of β and the changes in the income gap for the nine subperiods demonstrate a clear negative relationship between them. This suggests that temporal fluctuations of the catch-up process elicited by Barro type regression, when applied to short periods, qualitatively correspond to short-run dynamics across prefectures by the PANIC method.

In summary, section 4 reveals that the PANIC method is an effective vehicle that accurately draws a high-resolution picture of the long and short-run dynamics of Japanese regional economies.

5 Characterizing steady state real income

5.1 Predictions based on estimation results of the previous section

In the previous section, we characterized the long-run and short-run dynamic processes of prefectural per capita income. What prediction of the infinite future of income gap between prefectures can be drawn from the analysis? In the previous section, we find that permanent cross prefecture differences μ_i and the coefficient of the common factor λ_i are negatively correlated. Therefore, if the common factor F_t truly has a unit root with a positive time trend, eventually follower-prefectures will become richer than leader-; and the reversed gap between them will rapidly expand, becoming infinitely large in the far future. However, the prediction that currently poor prefectures will exceed rich ones in the future is quite novel, and needs critical examination. The following subsections explore cross-sectional properties of steady state per capita income using the framework proposed by Phillips and Sul (2007).

[23]The average annual speed of convergence over the period 1955–1999 is 2.1%, which is consistent with empirical evidence from earlier literature that reported very low convergence speeds (e.g. Barro and Sala-i-Martin, 1992b, report about 2%).

5.2 Relative convergence

The specification in the PANIC method that y_{it} linearly depends on F_t is inevitable for estimating F_t by principal component analysis. Bai and Ng (2004)'s approach of splitting y_{it} into μ_i, F_t, and e_{it} powerfully describes the situation as shown in the previous section. The specification of linear dependency of y_{it} on F_t is inappropriate, however, for analysis of the infinite future because the finding that F_t has a unit root under the linearity assumption predicts that the gaps between prefectures infinitely diverge. Although the forecast of prefectural per capita income in the infinite future critically depends on whether the process is a trend stationary or a unit root process, it is difficult to distinguish between stationary trend processes with a large autoregressive root and unit root processes in a finite sample.[24] We may incorrectly predict the income gap in the infinite future if we impose the restriction that a common component F_t has a unit root process without adequate examination.

In order to examine the possibility that F_t does not grow to infinity, we rewrite equation (6) following Phillips and Sul (2007), who proposed the concept of relative convergence. Their concept embodies the idea that economies converge if the cross-sectional *ratio* of dispersion of the Japanese real per capita income around a common trend decreases with time.

We rewrite equation (6) to allow for one idiosyncratic stochastic element by adding a time varying factor loading coefficient, as follows:

$$y_{it} = \delta_{it} F_t, \tag{9}$$

where

$$\delta_{it} \equiv \mu_i / F_t + \lambda_i + e_{it} / F_t \tag{10}$$

is a time varying prefecture-specific component that measures the time-varying economic distance between y_{it} and the common component F_t, and therefore we call it relative per capita income. If $\delta_{it} \to \delta_i$ for all i as $t \to \infty$, the relative real income between the i-th and

[24] Campbell and Perron (1991), Cochrane (1991), and Faust (1996) call this the near observational equivalence problem with unit root tests.

the j-th prefectures in the steady state is given as[25]

$$\lim_{t \to \infty} \frac{y_{it}}{y_{jt}} = \frac{\delta_i}{\delta_j}. \tag{11}$$

To calculate the steady state value of the time-varying relative per capita income of prefecture i, δ_{it}, we model its behavior as follows:

$$\delta_{it} = d_i + \phi_i \delta_{it-1} + \sum_{l=1}^{L} \phi_{il} \Delta\delta_{it-l} + \xi_{it}, \tag{12}$$

where d_i is a fixed parameter, ϕ_i is a prefecture-specific autoregressive parameter, and ξ_{it} is the error term with a zero mean. The lagged difference terms $\Delta\delta_{it-l}$ for $l = 1, \cdots, L$ are included to grasp a higher order serial correlation in the time series process for δ_{it}. The number of lags, L, is chosen to ensure that the remaining error terms ξ_{it} would be serially uncorrelated, resulting in lag four. For each prefecture, we calculate the value of δ_{it} from equation (9) and use these values to estimate equation (12) by ordinary least squares. The steady state of the time-varying relative per capita income δ_{it} is defined as $\delta_i \equiv d_i/(1-\phi_i)$. If F_t truly has a unit root, so that $F_t \to \infty$, as $t \to \infty$, then from equation (10) δ_i should equal λ_i. Thus, in the next subsection, we check whether δ_i equals λ_i to see if F_t truly has a unit root and therefore if F_∞ will reach infinity.

5.3 Although poor stay poor, the gap shrinks in the infinite future

We check the correlation between δ_i and λ_i, which is shown on the scatter plot in Figure 9. It appears that the steady state δ_i is unrelated to the fixed factor loading λ_i. Indeed, the coefficient of determination R^2 is 0.04. This suggests that the common component F_t does not truly have a unit root process, so that F_t does not go infinity as $t \to \infty$.

The result implies that the poor prefectures at the initial time will not infinitely dominate the rich ones. Then, the next question we address is if the poor will exceed the rich or not. We investigate the correlation between the steady state δ_i and its initial value δ_{i1956}. The scatter plot in Figure 10 reveals a statistically significant strong positive correlation ($R^2 = 0.45$). This implies that the relatively poorer prefectures at the initial

[25] Phillips and Sul (2007) point out that the relative real income in the steady state can be defined for any case regardless of whether F_t approaches infinity or not.

point (i.e., lower δ_{i1956}) will remain relatively poor in its steady state (i.e., lower δ_i). This phenomenon has been referred to as 苑 oor stay poor by Canova and Marcet (1995) and Shioji (2004).

The final question we address is whether the relative income gap will shrink or widen in the infinite future. To this end, we compare the Gini index of relative per capita income in the initial year δ_{i1956}, with that in the infinite future δ_i. The Gini indices calculated for δ_{i1956} and δ_i are 0.0175 and 0.0084, respectively, implying that the relative income gap between prefectures will eventually narrow. Thus, although "poor stay poor" in Japan, the poor prefectures catch up with rich ones substantially.

6 Concluding remarks

This paper analyzed regional growth dynamics in Japan using panel data of log per capita real income across Japanese prefectures. The Evans and Karras type of panel unit root test, which is powerful for evaluating the convergence hypothesis, did not adequately produce a detailed picture characterizing prefecture development. Thus, we examined the common trend hypothesis developed by Bernard and Durlauf (1995). The Johansen test was not able to investigate the cointegration relationship because of a limited degree of freedom, so we used the PANIC method provided by Bai and Ng (2004). Although the method described the long and short-run behaviors of regional growth well, it did not appropriately illustrate the infinite future. Thus, we employed the concept of relative convergence developed by Phillips and Sul (2007) to describe the steady state.

Our results are summarized by five main findings. First, applying the panel unit root test developed by Evans and Karras (1996), we found that the regional economies do not converge, implying that they do not have a single parallel balanced growth path.

Second, we adopted the PANIC method to make the cointegration approach applicable to an investigation of whether or not there are common sources of growth amongst the Japanese prefectures. We identified the common factor F_t, which was almost identical to the mean per capita income of Japan. Then, using unit root and panel cointegration tests, we discovered that the common component is nonstationary, and idiosyncratic components are stationary for all prefectures. This implies that Japanese prefectures have heterogeneous long-run balanced growth paths.

Third, the prefectures were classified as leader or follower-prefectures. The leader-prefectures have a larger permanent factor and smaller coefficient on the common trend than the follower-prefectures, implying that follower-prefectures will catch up to leader-prefectures.

Fourth, short-run deviations from the long-run equilibrium path indicate that the rate of shrinkage in income gaps among prefectures temporarily becomes slower or faster. The change in pace is synchronized with the transition of the Gini index of income gaps among prefectures.

Fifth, we used the Phillips and Sul (2007) method and found that the poor stay poor, but the relative income gap will narrow substantially in the infinite future.

Needless to say, there remain many problems to be solved. Among them, although we characterized the long-run and short-run income dynamics for each prefecture, where these characteristics for each prefecture originate is still an open question. This constitutes an interesting future task.

Acknowledgements

We are grateful to Daisuke Ishikawa, Ryuzo Miyao, Toshihiro Okubo and seminar participants at Osaka University, Miyazaki Sangyo-Keiei University, Toyama University, and the 2009 Japanese Economic Association Autumn Meeting at Senshu University, for their helpful comments and suggestions.

References

Bai, Jushan, Serena Ng. 2002 "Determining the Number of Factors in Approximate Factor Models", *Econometrica*, **70**(1): 191–221.

Bai, Jushan, Serena Ng. 2004 "A PANIC Attack on Unit Roots and Cointegration", *Econometrica*, **72**(4): 1127–1177.

Barro, Robert J., Xavier Sala-i-Martin. 1992a "Convergence", *Journal of Political Economy*, **100**(2): 223–251.

Barro, Robert J., Xavier Sala-i-Martin. 1992b "Regional Growth and Migration: A Japan-United States Comparison", *Journal of the Japanese and International Economies*, **6**(4): 312–346.

Barro, Robert J.. 1991 "Economic Growth in a Cross Section of Countries", *Quarterly Journal of Economics*, **106**(2): 407–443.

Bernard, Andrew B., Steven N. Durlauf. 1995 "Convergence in International Output", *Journal of Applied Econometrics*, **10**(2): 97–108.

Campbell, John Y., Pierre Perron. 1991 "Pitfalls and Opportunities: What Macroeconomists Should Know About Unit Roots", In Olivier Jean Blanchard, Stanley Fischer eds. *NBER Macroeconomics Annual*, Cambridge, MA. MIT Press: 141–220.

Canova, Fabio, Albert Marcet. 1995 "The Poor Stay Poor: Non-Convergence Across Countries and Regions", *CEPR Discussion Papers*. No. 1265.

Choi, In. 2001 "Unit Root Tests for Panel Data", *Journal of International Money and Finance*, **20**(2): 249–272.

Cochrane, John H.. 1991 "A Critique of the Application of Unit Root Tests", *Journal of Economic Dynamics and Control*, **15**(2): 275–284.

Dickey, David A., Wayne A. Fuller. 1979 "Distribution of the Estimators for Autoregressive Time Series With a Unit Root", *Journal of the American Statistical Association*, **74**(366): 427–431.

Durlauf, Steven N., Paul A. Johnson, Jonathan R.W. Temple. 2005 "Growth Econometrics", In Philippe Aghion, Steven N. Durlauf eds. *Handbook of Economic Growth*, **1**. North-Holland: 555–677.

Elliott, Graham, Thomas J. Rothenberg, James H. Stock. 1996 "Efficient Tests for an Autoregressive Unit Root", *Econometrica*, **64**(4): 813–836.

Evans, Paul, Georgios Karras. 1996 "Convergence Revisited", *Journal of Monetary Economics*, **37**(2): 249–265.

Faust, Jon. 1996 “Near Observational Equivalence and Theoretical Size Problems with Unit Root Tests”, *Econometric Theory*, **12**(4): 724–731.

Islam, Nazrul. 1995 “Growth Empirics: A Panel Data Approach”, *Quarterly Journal of Economics*, **110**(4): 1127–1170.

Islam, Nazrul. 1998 “Growth Empirics: A Panel Data Approach–A Reply”, *Quarterly Journal of Economics*, **113**(1): 325–329.

Johansen, Soren. 1995 *Likelihood-Based Inference in Cointegrated Vector Autoregressive Models*, Oxford. Oxford University Press.

Kawagoe, Masaaki. 1999 “Regional Dynamics in Japan: A Reexamination of Barro Regressions”, *Journal of the Japanese and International Economies*, **13**(1): 61–72.

Kwiatkowski, Denis, Peter C. B. Phillips, Peter Schmidt, Yongcheol Shin. 1992 “Testing the Null Hypothesis of Stationarity Against the Alternative of a Unit Root: How Sure are We That Economic Time Series Have a Unit Root?”, *Journal of Econometrics*, **54**(1-3): 159–178.

Lee, Kevin, M. Hashem Pesaran, Ronald Smith. 1997 “Growth and Convergence in Multi-Country Empirical Stochastic Solow Model”, *Journal of Applied Econometrics*, **12**(4): 357–392.

Lee, Kevin, M. Hashem Pesaran, Ronald Smith. 1998 “Growth Empirics: A Panel Data Approach–A Comment”, *Quarterly Journal of Economics*, **113**(1): 319–323.

Pesaran, Hashem M.. 2007 “A Pair-Wise Approach to Testing for Output and Growth Convergence”, *Journal of Econometrics*, **138**(1): 312–355.

Phillips, Peter C. B., Sam Ouliaris. 1990 “Asymptotic Properties of Residual Based Tests for Cointegration”, *Econometrica*, **58**(1): 165–193.

Phillips, Peter C. B., Donggyu Sul. 2007 “Transition Modeling and Econometric Convergence Tests”, *Econometrica*, **75**(6): 1771–1855.

Quah, Danny T.. 1996 “Empirics for Economic Growth and Convergence”, *European Economic Review*, **40**(6): 1353–1375.

Shibamoto, Masahiko, Yoshiro Tsutsui. 2011 "Note on the Interpretation of Convergence Speed in the Dynamic Panel Model", RIEB Discussion Paper Series DP2011-04, Kobe University.

Shin, Yongcheol. 1994 "A Residual-Based Test of the Null of Cointegration Against the Alternative of No Cointegration", *Econometric Theory*, **10**(1): 95–115.

Shioji, Etsuro. 2001 "Public Capital and Economic Growth: A Convergence Approach", *Journal of Economic Growth*, **6**(3): 205–227.

Shioji, Etsuro. 2004 "Initial Values and Income Convergence: Do "The Poor Stay Poor"?", *Review of Economics and Statistics*, **86**(1): 444–446.

Westerlund, Joakim, David L. Edgerton, Sonja Opper. 2010 "Why is Chinese Provincial Output Diverging?", *Journal of Asian Economics*, **21**(4): 333–344.

Yamane, Chisako, Yoshiro Tsutsui. 2009 "Banking Sector and Regional Economic Growth", *Review of Monetary and Financial Studies*, **28**(1): 68–90. (In Japanese).

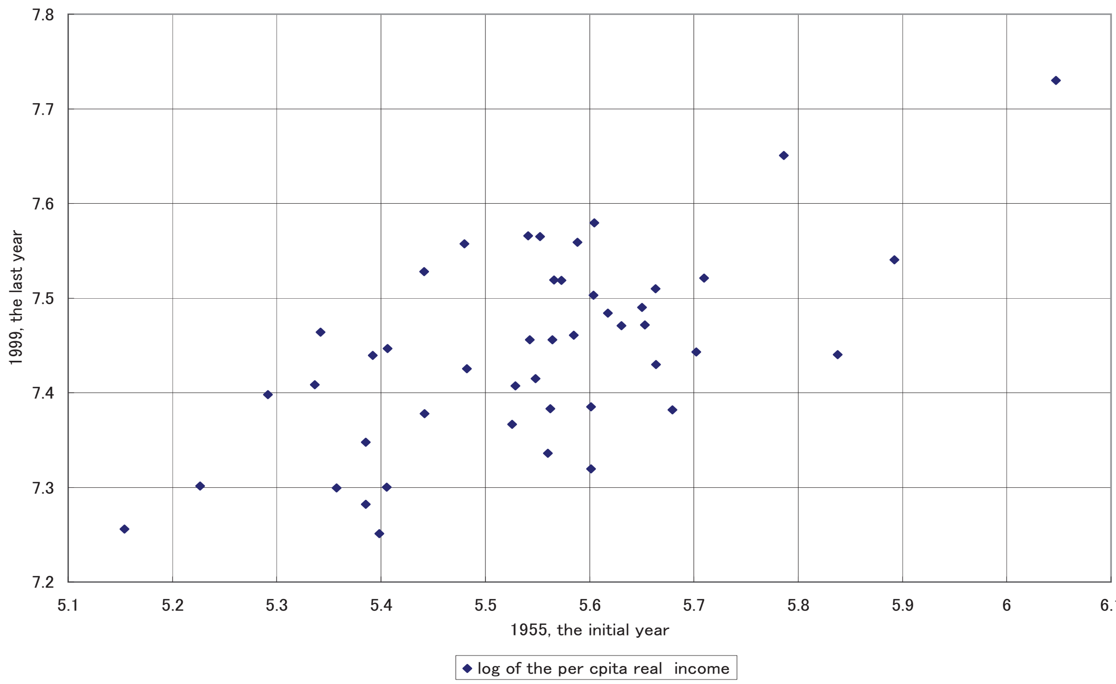

Figure 1: Scatter plot of the log per capita real income, 1955 vs. 1999

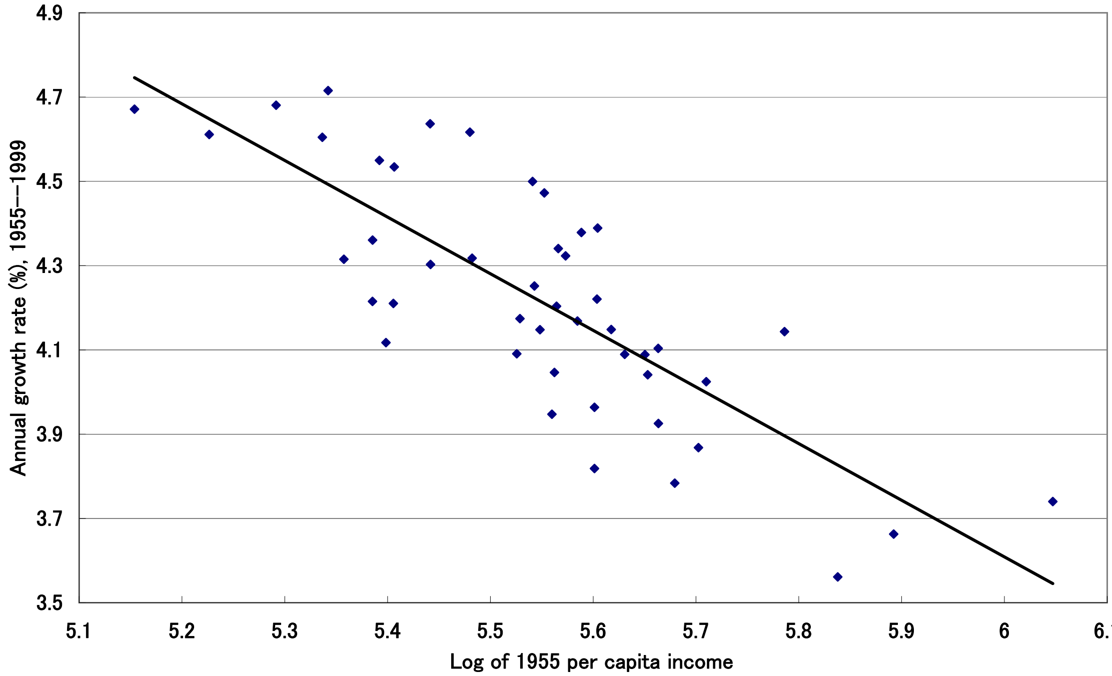

Figure 2: Scatter plot of the per capita income in 1955 vs. income growth in 1955–1999 across prefectures

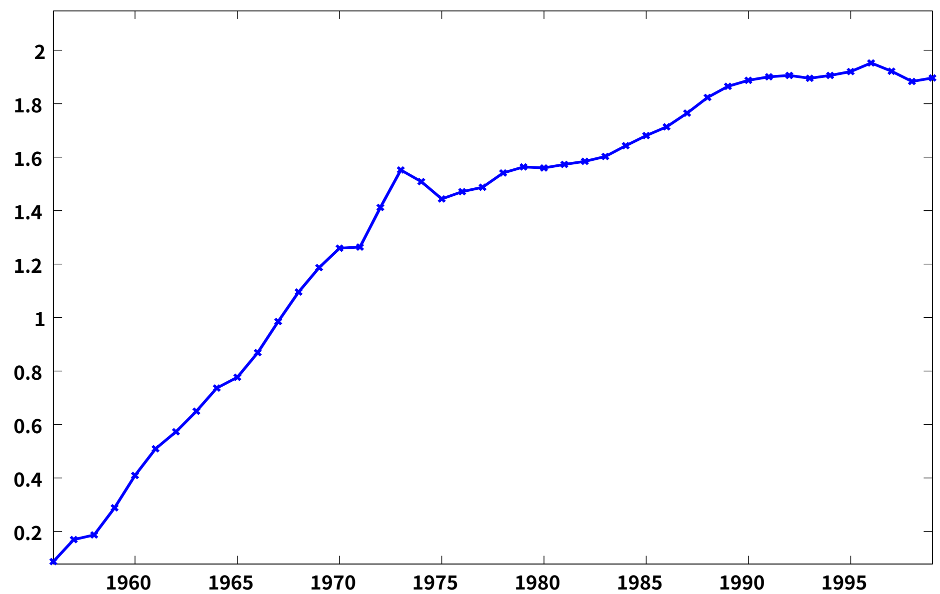

Figure 3: Common component

Note: The common component is determined by applying the principal component analysis to equation (7) in the text and calculating the integrated series of the principal component. The Bai and Ng (2002)'s information criteria indicates that only a single common component exists.

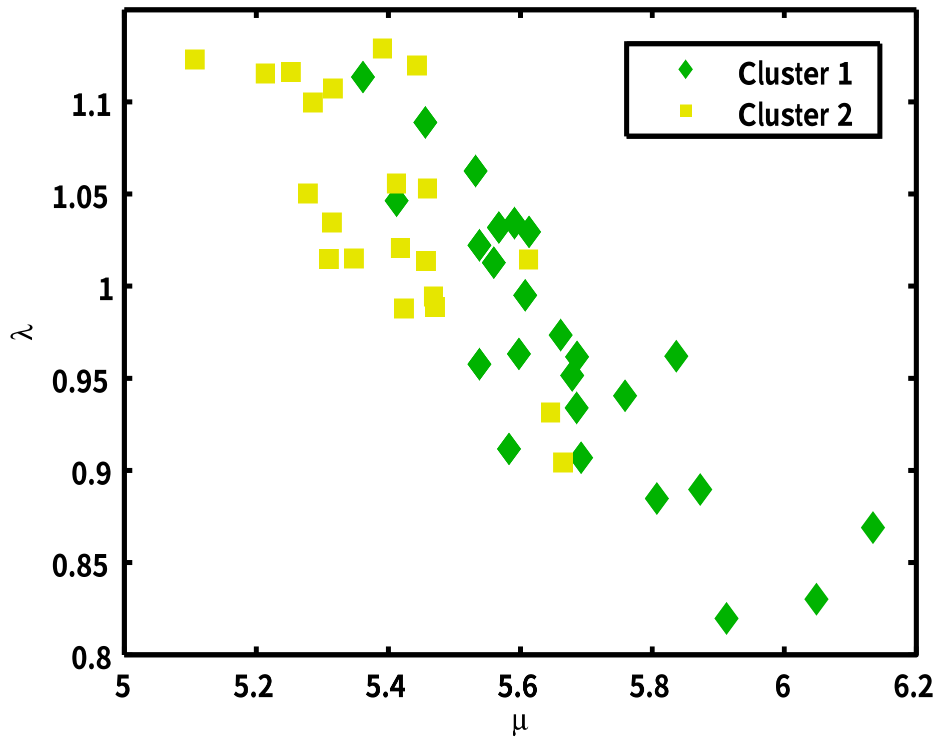

Figure 4: Scatter plot of permanent difference μ_i vs. weight to a common component λ_i across prefectures

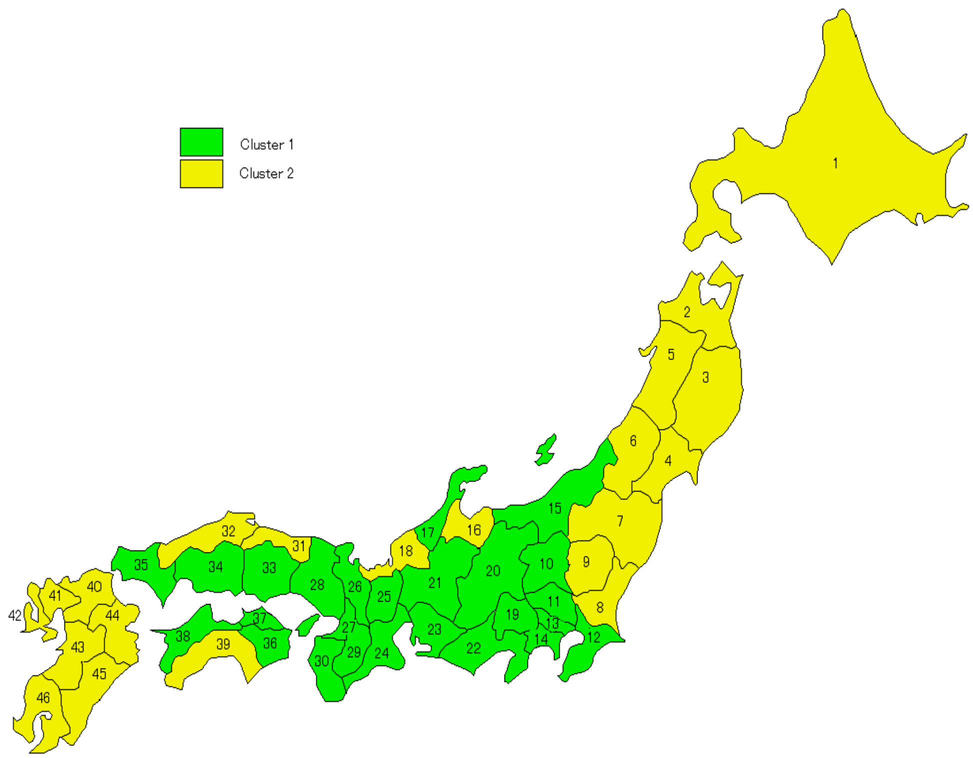

Figure 5: Two clusters of prefectures stratified by the estimates of prefecture-specific components $\hat{e}_{it}$

Note: In this figure, the prefectures of Japan are colored according to clusters as explained in the map legend. The prefectures where grouped into clusters by applying k-means cluster analysis to the prefecture-specific components $\hat{e}_{it}$ so as to maximize the similarities of the constituents. The numbers in the map refer to the prefecture indexes. The prefecture name of each index number is given in Table 2.

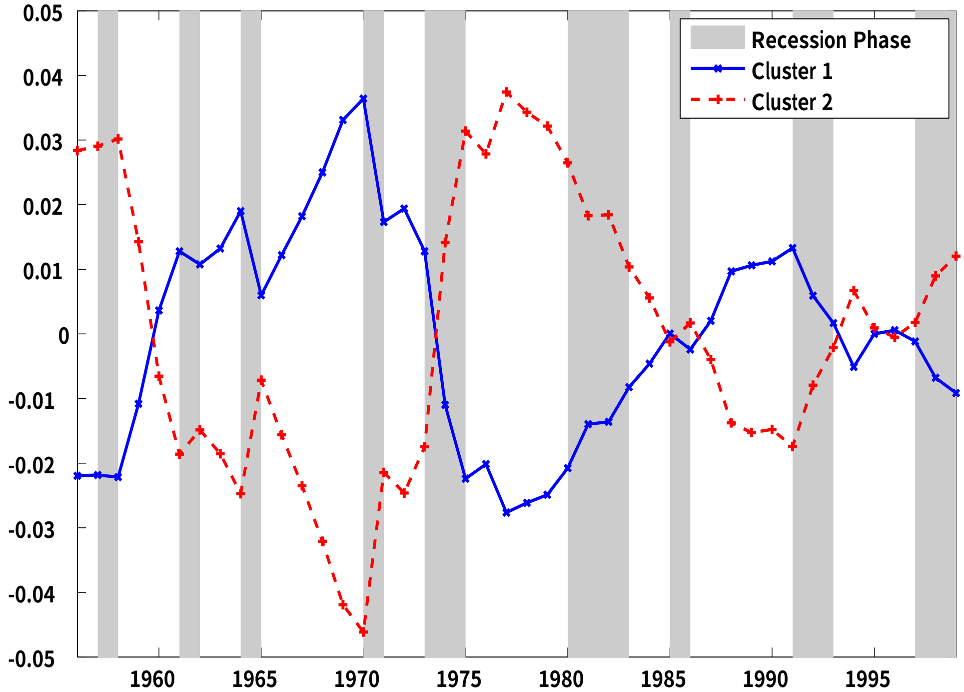

Figure 6: Short-run deviations e_{it} of clusters 1 and 2.

Note: The graphs of clusters 1 and 2 are symmetrical by construction, since the total deviation from the average is zero. The shaded bars in the background indicate recession periods.

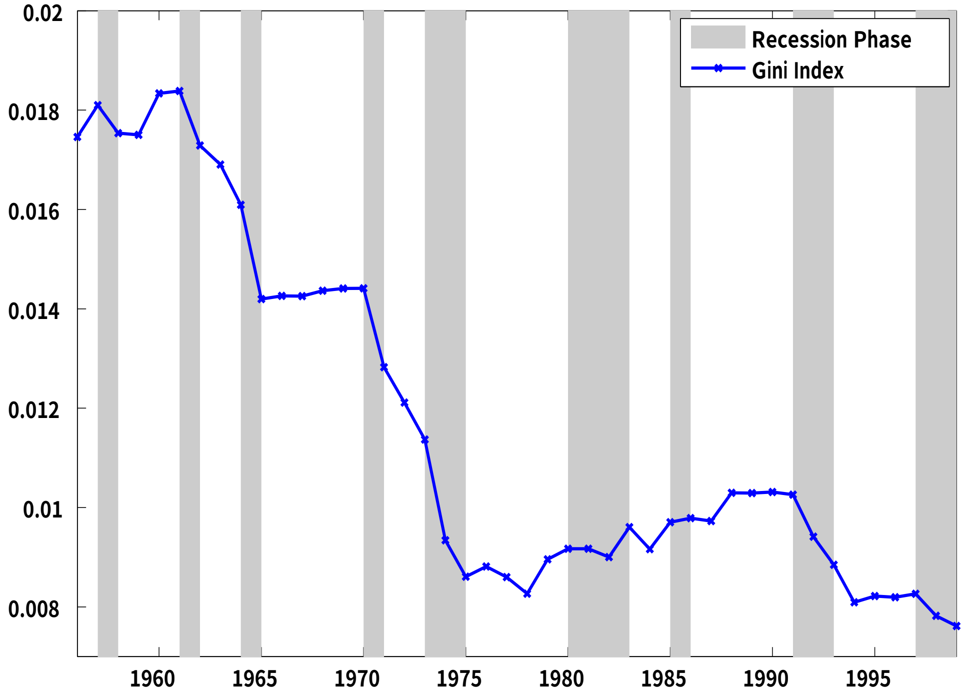

Figure 7: Gini index of the prefectural per capita income

Note: The shaded bars in the background indicate recession periods.

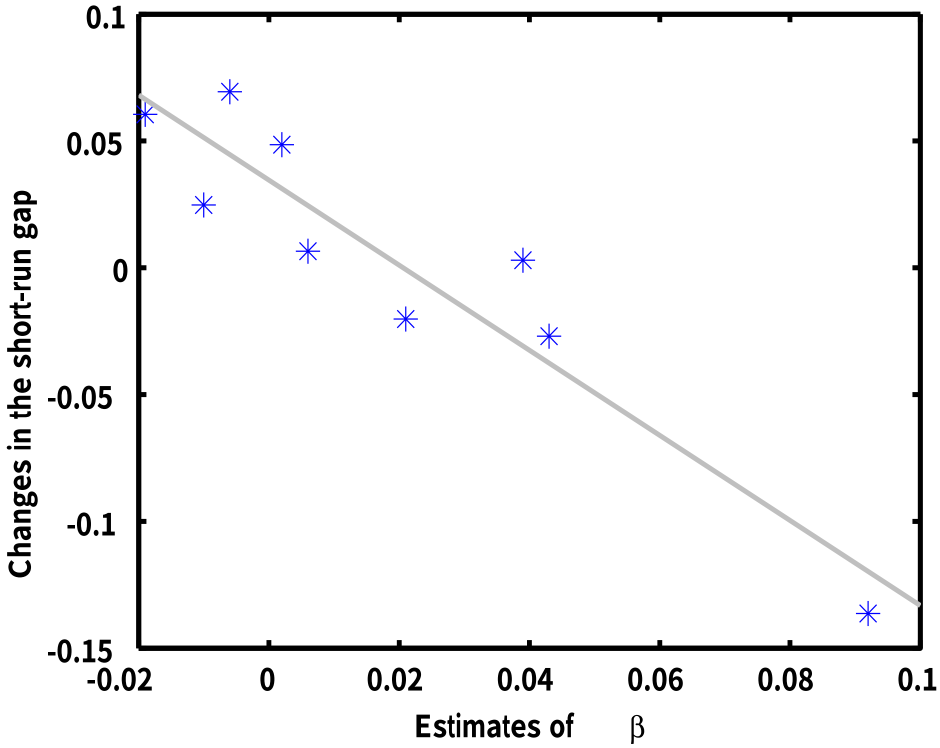

Figure 8: Scatter plot of the convergence speed vs. the changes in the short-run gap between clusters 1 and 2 for nine subperiods

Note: Each subperiod is five years long starting from 1955. The fitted line to the nine observations is shown in the figure.

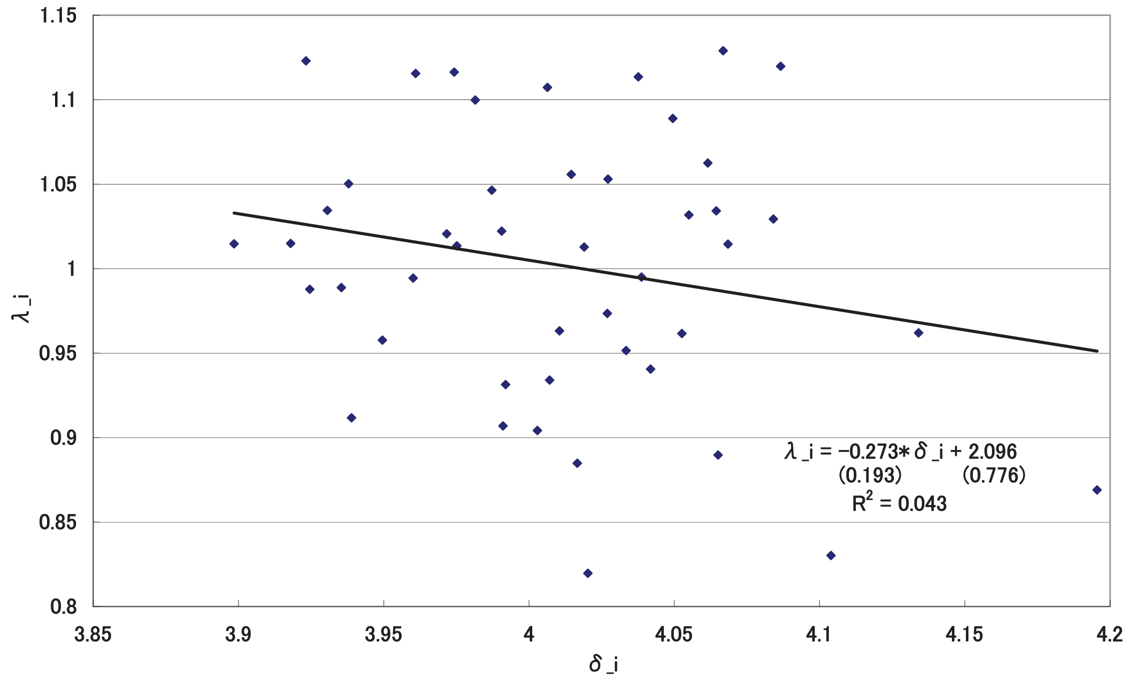

Figure 9: Scatter plot of relative income δ_i vs. the weight to a common component λ_i in the steady state across prefectures

Note: Fitted line and the estimation results of the regression are shown in the graph. Standard errors are given in parentheses.

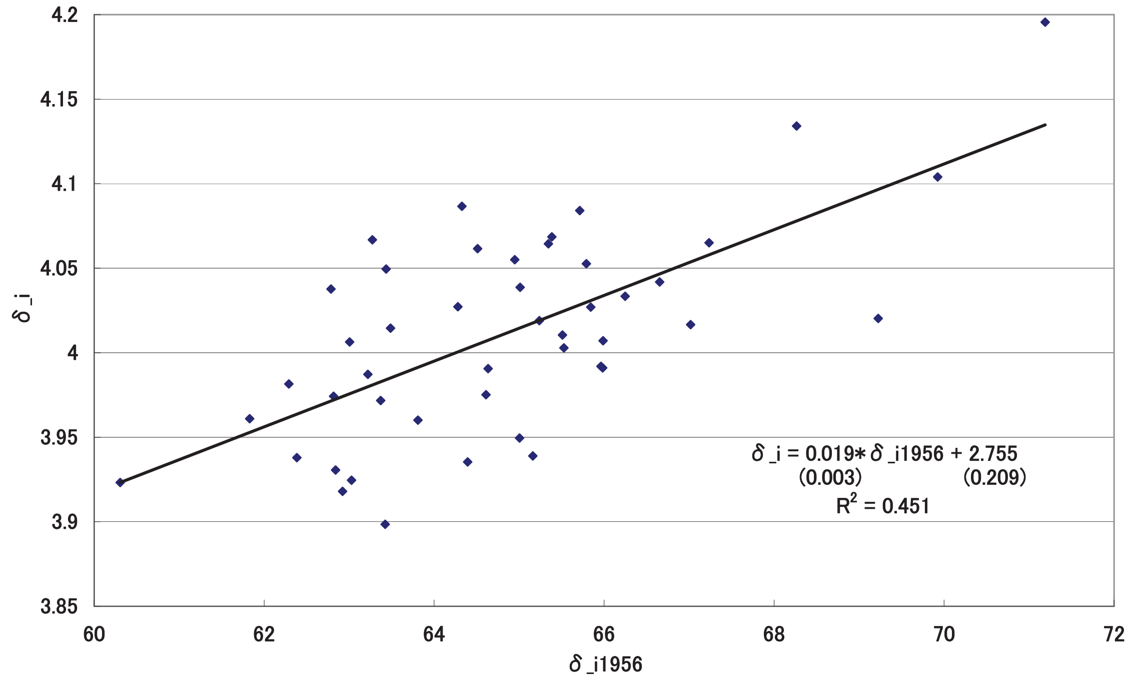

Figure 10: Scatter plot of the initial and steady state value of the relative income δ_i across prefectures

Note: The fitted line and the estimation results of the regression are shown in the graph. Standard errors are given in parentheses.

Table 1: Panel unit root test: Evans and Karras (1996)'s type convergence test

Variable	ADF	KPSS	$\hat{\rho}$	F stat
$y_{it} - \bar{y}_t$	2.182^{**}	2.209^{**}	0.88^{***}	1.40^{***}

Note: ADF is a Fisher's statistic P_N as defined in Choi (2001) based on a p value of the individual augmented Dickey and Fuller (1979) test of the null of a unit root. The lag length of the lagged difference terms, $y_{it} - \bar{y}_t$, to be added to the individual ADF test was set at 4. KPSS is a Fisher's statistic P_N as defined in Choi (2001) based on a p value of the individual Kwiatkowski et al. (1992) test of the null of no unit root. The number of truncation lags in the KPSS test was set at 12. A Fisher's statistic P_N as defined in Choi (2001) has a $N(0,1)$ distribution under the null hypothesis. $\hat{\rho}$ is the pooled estimate of the autoregressive parameter ρ_i of $y_{it} - \bar{y}_t$ in equation (3). F stat stands for the F statistic under the null that the autoregressive parameter ρ_i is identical for all $i = 1, \cdots, N$. *** and ** denote rejection of the null hypothesis by the 1% and 5% significance levels, respectively.

Table 2: Individual unit root test: Evans and Karras (1996)'s convergence test

Index	Name	ADF	DFGLS	KPSS	$\hat{\rho}_i$	Index	Name	ADF	DFGLS	KPSS	$\hat{\rho}_i$
1	Hokkaido	−2.90**	−1.13	0.499**	0.80	24	Mie	−1.28	−1.32	0.300	0.88
2	Aomori	−1.69	−1.73*	0.445*	0.76	25	Shiga	−1.13	−0.72	0.297	0.89
3	Iwate	−0.66	0.22	0.586**	0.97	26	Kyoto	−1.28	−0.47	0.509**	0.94
4	Miyagi	−1.31	−1.37	0.486**	0.89	27	Osaka	−0.29	−0.18	0.587**	0.99
5	Akita	−3.62***	−2.19**	0.211	0.36	28	Hyogo	−3.24**	−0.31	0.564**	0.86
6	Yamagata	−2.49	−3.05***	0.407*	0.52	29	Nara	−2.77*	−0.68	0.411*	0.83
7	Fukushima	−0.67	−0.56	0.513**	0.96	30	Wakayama	−1.05	−0.44	0.549**	0.90
8	Ibaraki	−1.33	−0.90	0.55**	0.92	31	Tottori	−2.63*	−1.20	0.092	0.67
9	Tochigi	−1.77	−1.05	0.505**	0.90	32	Shimane	−2.69*	−2.30**	0.206	0.62
10	Gunma	−2.76*	−1.36	0.476**	0.85	33	Okayama	−2.43	−1.56	0.138	0.71
11	Saitama	−2.37	−1.33	0.393*	0.78	34	Hiroshima	−1.31	−1.05	0.525**	0.92
12	Chiba	−2.31	−1.15	0.327	0.82	35	Yamaguchi	−1.74	−1.58	0.391*	0.86
13	Tokyo	−1.25	−0.53	0.519**	0.94	36	Tokushima	−0.40	−0.84	0.591**	0.96
14	Kanagawa	−0.96	−0.82	0.487**	0.95	37	Kagawa	−2.58*	−2.10**	0.108	0.75
15	Niigata	−1.83	−1.43	0.233	0.84	38	Ehime	−1.68	−0.92	0.51**	0.78
16	Toyama	−2.50	−2.67***	0.144	0.64	39	Kochi	−1.49	−1.46	0.185	0.89
17	Ishikawa	−2.78*	−2.84***	0.125	0.59	40	Fukuoka	−2.37	−1.08	0.453*	0.83
18	Fukui	−1.47	−1.20	0.381*	0.91	41	Saga	−2.16	−1.69*	0.093	0.67
19	Yamanashi	−1.84	−0.80	0.544**	0.88	42	Nagasaki	−2.49	−1.72*	0.083	0.70
20	Nagano	−4.92***	−1.90*	0.243	0.48	43	Kumamoto	−2.49	−1.44	0.439*	0.90
21	Gifu	−2.13	−1.44	0.354*	0.84	44	Oita	−0.33	−0.79	0.543**	0.98
22	Shizuoka	−2.50	−1.90*	0.369*	0.74	45	Miyazaki	−2.55*	−0.64	0.328	0.70
23	Aichi	−2.43	−1.78*	0.193	0.79	46	Kagoshima	−1.22	−0.66	0.482**	0.94

Note: ADF is the augmented Dickey and Fuller (1979) test of the null of a unit root. DFGLS is a Dickey-Fuller test based on GLS-detrended series, proposed by Elliott et al. (1996). The lag length of the lagged difference terms, $y_{it} - \bar{y}_t$, to be added to the ADF test and DFGLS test was set at 4. KPSS is the Kwiatkowski et al. (1992) test of the null of no unit root. The number of truncation lags in the KPSS test was set at 12. We used the critical values presented in Dickey and Fuller (1979), Elliott et al. (1996), and Kwiatkowski et al. (1992). ***, **, and * denote rejection of the null hypothesis by the 1%, 5%, and 10% significance levels, respectively. $\hat{\rho}_i$ is the estimate of the autoregressive parameter ρ_i of $y_{it} - \bar{y}_t$ in equation (3).

Table 3: Unit root and cointegration tests for common and idiosyncratic components

Common component F_t	
ADF	KPSS
-1.748	0.158^{**}

Idiosyncratic components $\hat{e}_{it}$	
PO	Shin
2.548^{**}	-0.581

Note: Unit root tests for the common trend F_t include a constant and a linear trend. As for ADF and KPSS, see the footnote of Table 2. PO is a Fisher's statistic defined as in Choi (2001) P_N based on a p value of the individual Phillips and Ouliaris (1990) test, where the null hypothesis is no cointegration. The lag length of the lagged difference terms of $\hat{e}_{it}$ to be added to individual Phillips and Ouliaris (1990) test was set at 4. Shin is a Fisher's type statistic defined as in Choi (2001) P_N based on a p value of the individual Shin (1994) test where the null hypothesis is cointegration. The number of truncation lags in the individual Shin (1994) test was set at 12. A Fisher's statistic P_N as defined in Choi (2001), has a $N(0, 1)$ distribution under the null hypothesis. *** and ** denote rejection of the null hypothesis by the 1% and 5% significance levels, respectively.

Table 4: Nonlinear regressions (8) for per capita real income across Japanese prefectures

Period	$\hat{\beta}$	adj-R^2	Period	$\hat{\beta}$	adj-R^2
1955–1999	0.021***	0.623	1975–1980	0.006	−0.015
	(0.008)			(0.007)	
1955–1960	−0.019**	0.060	1980–1985	0.002	−0.022
	(0.008)			(0.011)	
1960–1965	0.039***	0.430	1985–1990	−0.010	0.021
	(0.007)			(0.006)	
1965–1970	−0.006	−0.005	1990–1995	0.043***	0.600
	(0.007)			(0.007)	
1970–1975	0.092***	0.685	1995–1999	0.021***	0.190
	(0.015)			(0.005)	

Note: Figures in parentheses are heteroskedasticity robust standard errors. *** and ** denote rejection of the null hypothesis by the 1% and 5% significance levels, respectively. The estimate of the constant term is suppressed.