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Masahiko SHIBAMOTO
Yoshiro TSUTSUI
Chisako YAMANE

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Research Institute for Economics and Business Administration

Kobe University

2-1 Rokkodai, Nada, Kobe 657-8501 JAPAN

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MASAHIKO SHIBAMOTO[a]

YOSHIRO TSUTSUI[b]

CHISAKO YAMANE[c]

[a] *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.

[b] Corresponding author: *Faculty of Economics, Konan University*,

8–9–1 Okamoto, Higashinada, Kobe, 658–8501 Japan. E–mail: tsutsui@center.konan-u.ac.jp.

[c] *Faculty of Economics, Okayama Shoka University*,

2–10–1 Tsushima, Kyomachi, Kitaku, Okayama, 700–8601 Japan. E–mail: yamane@po.osu.ac.jp.

SUMMARY

This study proposes a panel cointegration approach using the PANIC method for understanding the regional growth dynamics using nonstationary panel data, and applies it to Japanese prefectures. This approach enables us to analyze both long-run equilibrium growth path and short-run dynamics across the regions. Specifically, we find that there is one common source of growth to which prefectures attach different weights, that the per capita real income of follower-prefectures will catch up to that of leader-prefectures, and that temporal fluctuations of the catch-up process elicited by Barro type regression qualitatively corresponds to short-run dynamics across prefectures by the PANIC method.
JEL Classification: O40; C22; C23.

Keywords: β-convergence, common trends, panel unit root test, PANIC method, Japan.

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, that 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 time points 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 time points and tells us nothing about the dynamic process of the 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 (1995)'s Definition 2.1 and Evans and Karras (1996) provide a

[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 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".

test for absolute and conditional convergence hypotheses using a unit root test on panel data (hereinafter, the Evans and Karras type test). Their idea is that two economies will converge if the difference in per capita income between the two economies is stationary. This condition 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 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. There exists the possibility 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 occasion 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]

[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

Its idea is to estimate common factors of per capita incomes of the economies before estimating the cointegration relation between per capita income and the factors.[5]

We apply the method to panel data from 46 Japanese prefectures from 1955 to 1999. As shown in section 2, the income gap among prefectures in Japan has been evidently decreasing in a long-run perspective over the entire sample period, suggesting that the Japanese economy serves as a good exercise for the convergence problem. On the other hand, as discussed in subsection 4.5, the rate of change in the income gap between leader and follower prefectures has been fluctuating over the same period. We analyze the regional income dynamics underlying these findings.

Several novel features of our analysis set this study apart from the existing literature. First, we explore the long-run equilibrium growth across Japanese prefectures. Literature using panel data has not reached consensus on income convergence across them.[6] While panel unit root tests have generally accepted the no convergence null (e.g., Kawagoe, 1999), tests based on the dynamic panel regression approach, 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.

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

gaps across economies. This method may be an alternative to the method adopted in this paper. His approach is applicable when the number of the economies is large relative to the time dimension of the panel.

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

[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 pro and Bernard and Durlauf (1995) and Quah (1996) for con.

[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. A problem of this approach is argued in Shibamoto and Tsutsui (2014).

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. 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 are 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, total prefectural income is 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-war period" average of 1934–1936 is normalized at 1.

Figure 1 show a scatter plot of the log per capita income in the initial (1955) and final (1999) years of the study. The plot indicates a positive relationship between per capita income in 1955 and 1999. This implies that poor prefectures remain poor and rich prefectures remained rich.

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 in the

[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 definition of the income and outlay account and its calculation methods. Therefore, to avoid bias, we analyzed the period when the 68SNA data cover, as a benchmark. However, we set the ending point of the sample in 2012, including the period from 2000 to 2012 by using 93SNA data, and obtained qualitatively the same results as the benchmark ones. These results are available from the authors upon request.

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

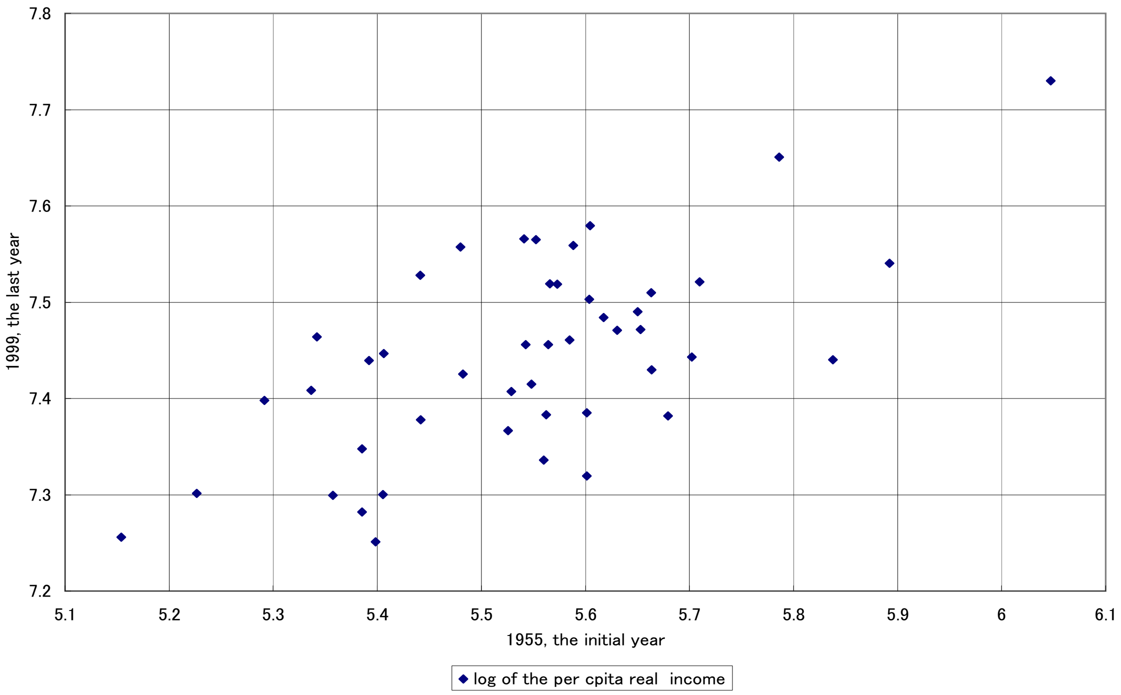

Figure 1. Scatter plot of the log per capita real income, 1955 vs. 1999

figure, i.e. the Barro regression, shows a strong negative correlation between them.

These two preliminary analyses indicate that, although follower-prefectures stay poor and leader-prefectures stay rich, 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, 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.

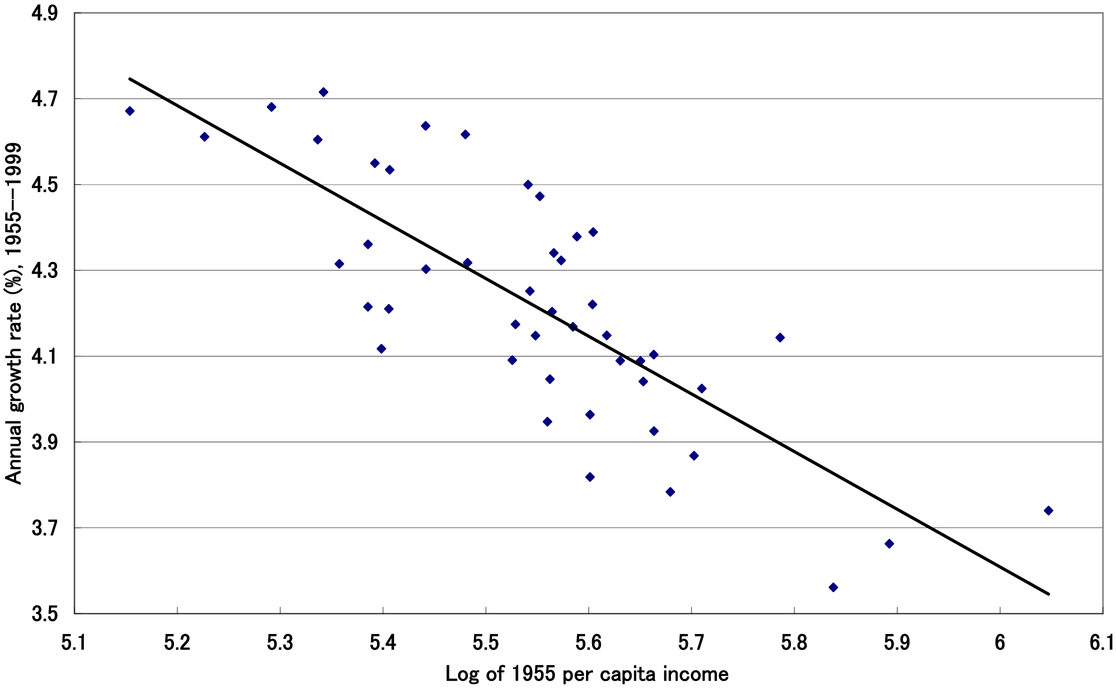

Figure 2. Scatter plot of the per capita income in 1955 vs. income growth in 1955–1999 across prefectures

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

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

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 the pooled tests based on Fisher's type statistics defined as in Choi (2001). The main idea for Choi (2001)'s tests is to combine p-values from a unit root/cointegration test applied to each economy under the null hypothesis, under which it is assumed that all cross section units have a unit root (or are stationary), against the alternative, under which it is assumed that some of all panel units are stationary (or have a unit root). Choi (2001)'s test statistics, which we call P_N statistics, is as follows:

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

where p_i is the p-value associated with unit root/cointegration test statistics for economy 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 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 Type Convergence Test

The second column of Table I presents a Choi (2001) statistics P_N based on the ADF type unit root test for the individual prefectures.[12] This allows us to test the null hypothesis that

[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.

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.

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

Variable	ADF	KPSS	$\hat{\rho}$	F stat
$y_{it} - \bar{y}_t$	2.161^{**}	12.450^{***}	0.88^{***}	1.40^{***}

Notes: ADF is a Fisher's type statistic P_N defined as 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 type statistic P_N defined as 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 type statistic P_N defined as in Choi (2001) has a $\mathbb{N}(0, 1)$ 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 statistics under the null that the autoregressive parameter ρ_i is identical for all $i = 1, \cdots, N$. *** and ** denote rejection of the null by the 1%, and 5% significance levels, respectively.

The third column of Table I presents a Choi (2001) statistics 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 fraction has 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 (2001)'s P_N, which are presented in Table II. 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

[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 find that there are some pairs for which the differences in output are stationary with a constant mean, although there is less evidence of output

the former group converges and the latter group does not, because measuring deviations from the average prefectural income $\bar{y}_t$ is questionable for this interpretation. 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 of identical ρ is rejected at the 1% significance level, as shown in the fifth column of Table I.

In sum, 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 sate 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. Although we have shown that Japanese prefectures do not have parallel balanced growth paths, they may have 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

convergence at a global level.

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

Table II. Individual unit root test: Evans and Karras (1996)'s type 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

Notes: 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 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).

Bernard and Durlauf (1995)'s 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. Our method is new, however, in the 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 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 (1995)'s 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 factors of incomes and then tests if the common factors 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 factor $\hat{F}_t$, 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, a constant and time trend are included.

To test the stationarity of an estimated idiosyncratic component $\hat{e}_{it}$, Bai and Ng (2004) propose using pooled tests based on the 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 tested 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 almost identically with 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).

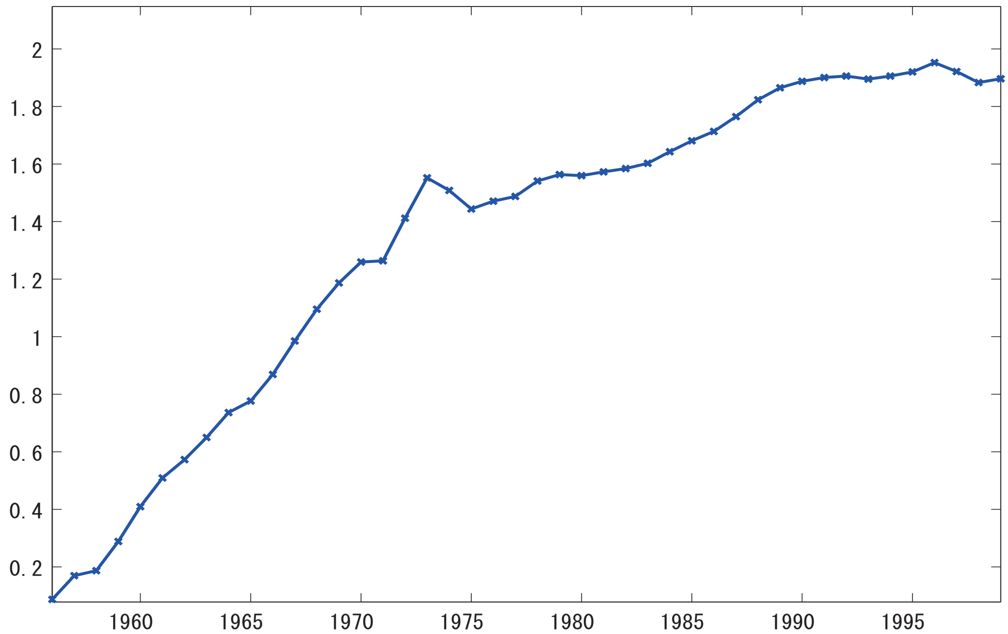

Figure 3. Common factor

Notes. The common factor 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 there exists only one common factor.

Next, we investigate the time-series properties of the common factor F_t and the prefecture-

[16]There might be a *significant* cross prefectural correlation explaining the additional common factor F_t, and then a violation of Assumption C(iii) from Bai and Ng (2004): e_{it} values are allowed to exhibit only *weak* cross sectional correlation. So in this paper we follow the approach of Bai and Ng (2004) by *statistically* determining the number of common factors r in equation (7) using the Bai and Ng (2002)'s information criteria. We adopted one of their six criteria, the most robust criterion when cross correlation is found 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 $IC(r)$.

specific component $\hat{e}_{it}$ for all prefectures. The upper panel of Table III presents the unit root test results of the common factor F_t. We find that the ADF test accepts the null that the common factor F_t has a unit root (the left column) and the KPSS test rejects the null 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.

Table III. Unit root and cointegration tests for common and idiosyncratic components

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

Idiosyncratic components $\hat{e}_{it}$	
PO	Shin
2.367^{**}	1.015

Notes: 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 II. PO is a Fisher's type 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 type statistic P_N defined as in Choi (2001) P_N has a $\mathbb{N}(0,1)$ distribution under the null hypothesis. *** and ** denote rejection of the null by the 1%, and 5% significance levels, respectively.

The lower panel of Table III presents the results on idiosyncratic components. We apply a Choi (2001)'s statistics P_N based on the Phillips and Ouliaris (1990) cointegration test to each prefecture. The null 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 that all the estimated prefecture-specific components $\hat{e}_{it}$ are stationary was accepted by a Choi (2001)'s statistics P_N based on the Shin (1994) cointegration test applied to each prefecture (the right column).[17]

Our findings in Table III indicate that prefectural per capita real incomes in Japan share

[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.

common sources of non-stationary variation in the form of a country-wide stochastic trend. This implies that the univariate nonstationary component characterizes the country-wide factors that cause changes in the economic growth of the respective prefectures, not necessarily just one source of growth in terms of the Solow model or its extended versions. As such, the common factor F_t and possible sources of the country-wide growth may share the same stochastic trend.

To statistically explore the characteristics of this stochastic trend, we examine whether there is any cointegration between the common factor F_t and country-wide growth factors. Based on a neoclassical growth model, we decompose real output per capita (denoted by Y/N) into the following multiple sources of growth;

$$\frac{Y}{N} = \frac{Y}{L}\frac{L}{N} = AK^{\phi}L^{-\psi}\frac{L}{N}, \tag{8}$$

where $Y = AK^{\phi}L^{1-\psi}$ (Cobb-Douglas production function), A denotes TFP, K denotes Real Capital, L denotes Labor Input, and N denotes Population. If we assume the constant returns to scale production function ($\phi = \psi$), labor productivity (Y/L) is decomposed into TFP (A) and Capital Intensity (K/L). If the stochastic trend in the common factor F_t imparts a stochastic trend to the economy-wide growth factors, their long-run relation is given by:

$$F_t = \delta^{\iota} + \delta^{t}t + \delta^{x\prime}x_t + \xi_t, \tag{9}$$

where x_t is (a vector of) the growth factors (log of Y/L, L/N, K/L, A, K, L, and their multivariate combinations) and ξ_t is a stationary error term. The stationary nature of ξ_t implies that the stochastic trends in the common factor F_t and the growth factors can be eliminated via the linear combination given by $\delta^{\iota}, \delta^{t}$ and δ^{x}, which is termed a cointegrating relationship.

Table IV presents the cointegration test results using 11 specifications for growth factors, as a variable of x_t: (1) Real GDP per worker (Y/L) and Labor-Population ratio (L/N); (2)

Real GDP per worker (Y/L); (3) TFP (A) and Real Capital Intensity (K/L); (4) TFP (A) and Real Capital (K); (5) TFP (A) and Worker (L); (6) Real Capital (K) and Worker (L); (7) Labor-Population ratio (L/N); (8) TFP (A); (9) Real Capital Intensity (K/L); (10) Real Capital (K); and (11) Worker (L).[18] The second column is a Phillips and Ouliaris (1990) test statistic, where the null hypothesis is no cointegration. The third column is a Shin (1994) test statistic where the null hypothesis is cointegration.

Table IV. Tests for cointegration between the common factor F_t and the growth factors

Growth factors	PO	Shin
(1) Real GDP per worker (Y/L), Labor-Population ratio (L/N)	-4.645^{***}	0.062
(2) Real GDP per worker (Y/L)	-4.211^{**}	0.089
(3) TFP (A), Real Capital Intensity (K/L)	-5.484^{***}	0.050
(4) TFP (A), Real Capital (K)	-5.270^{***}	0.050
(5) TFP (A), Worker (L)	-2.405	0.116^{**}
(6) Real Capital (K), Worker (L)	-2.585	0.085^{*}
(7) Labor-Population ratio (L/N)	-1.256	0.196^{***}
(8) TFP (A)	-2.433	0.122^{*}
(9) Real Capital Intensity (K/L)	-2.480	0.092
(10) Real Capital (K)	-2.513	0.093
(11) Worker (L)	-2.558	0.127^{**}

Notes: The tests for cointegration between common factor F_t and the growth factors include a constant and a linear trend. PO is a Phillips and Ouliaris (1990) test statistic, where the null hypothesis is no cointegration. The lag length of the lagged difference residuals ξ_t to be added to Phillips and Ouliaris (1990) test was set at 1. Shin is a Shin (1994) test statistic where the null hypothesis is cointegration. The number of truncation lags in the Shin (1994) test was set at 8. ***, ** and $*$ denote rejection of the null by the 1%, 5% and 10% significance levels, respectively.

In Table IV we find that the same stochastic trend in F_t is present in time series for the country-wide growth factors, especially labor productivity (Y/L) and the multivariates of TFP (A) and Capital Intensity(K/L) or Real Capital (K). First, we detect cointegrating relationship between the common factor F_t and possible underlying determinants of the real per capita output, Y/L and L/N, as reported in the first row of Table IV: both the Phillips

[18] The data used in this analysis are extracted from the Penn World Table 8.1 of Feenstra et al. (2015), as provided at www.ggdc.net/pwt and www.internationaldata.org. They consist of annual real GDP (rgdpna), TFP (rwtfpna), Real Capital (rkna), Labor input (emp), and Population (pop) in Japan over the period of 1955–2011.

and Ouliaris (1990) and Shin (1994) tests support a stationary residual from equation (9). In the second (third, respectively) row, we confirm the presence of cointegration between F_t and labor productivity Y/L (multivariate combination of TFP (A) and Real Capital Intensity (K/L)), which impylies that the Labor-Population Ratio, L/N, does not play a crucial role in the cointegration relation. In addition, the fourth row indicates the presence of cointegration between F_t and the combination of TFP (A) and Real Capital (K), while the fifth (sixth) row shows that there is no cointegration relationship between F_t and the combination of TFP (A) and Labor input (L) (Real Capital (K) and Labor input (L)). This implies that the linear combination of TFP (A) and Real Capital (K) plays an important role in characterizing the source of the stochastic trend in prefectural per capita real incomes. Finally, the test results in the seventh to eleventh rows do not support the presence of a cointegration relation between F_t and any of the growth factors. This implies that it would be difficult to characterize the stochastic trend in the common factor F_t by only one of the growth factors.

In sum, 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 and contains the stochastic trend characterized by the country-wide labor productivity or TFP and Real Capital. However, each prefecture attaches a different weight, λ_i, to it, implying that the prefectures follow different long-run balanced growth paths.[19]

4.4. Cross-Prefectural Properties of the Long-Run Equilibrium Growth Path

The results of Table III elucidate the reason why 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 $y_{it} - y_{jt} =$

[19]Note that this result does not imply that the β-convergence hypothesis holds. On the other hand, researchers using the dynamic panel regression approach, e.g., Lee et al. (1997, 1998), argue that economies converge at a relatively high β convergence speed compared to the analyses of the conventional cross-sectional regression. We argue that their estimates do not reflect the convergence speed in the long-run, but the adjustment speed of the short-run deviation from the long-run equilibrium path. See Shibamoto and Tsutsui (2014) for more details of this interpretation.

$\mu_i - \mu_j + (\lambda_i - \lambda_j)F_t + e_{it} - e_{jt}$. Given that F_t is a 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 the others are not, really means that the factor loading coefficients coincide for some prefectures, but not for others. This argument is similar to Pesaran (2007).[20]

The long-run equilibrium is characterized with 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 effect. 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 empirically shown in Section 3.

4.5. Cross-Prefectural Properties of the Short-Run Dynamic Behavior

Next we examine cross-prefectural properties of the short-run dynamic behavior. Since it is difficult to draw clear conclusions from the examination of each pattern of 46 prefectures, which we have done preliminarily, 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.[21] We point out the following five facts.

First, in Figure 5, map of Japan is painted over with different brightness which shows

[20] 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 (1995) Definition 2.1, if $\lambda_i = \lambda_j$.

[21] 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.

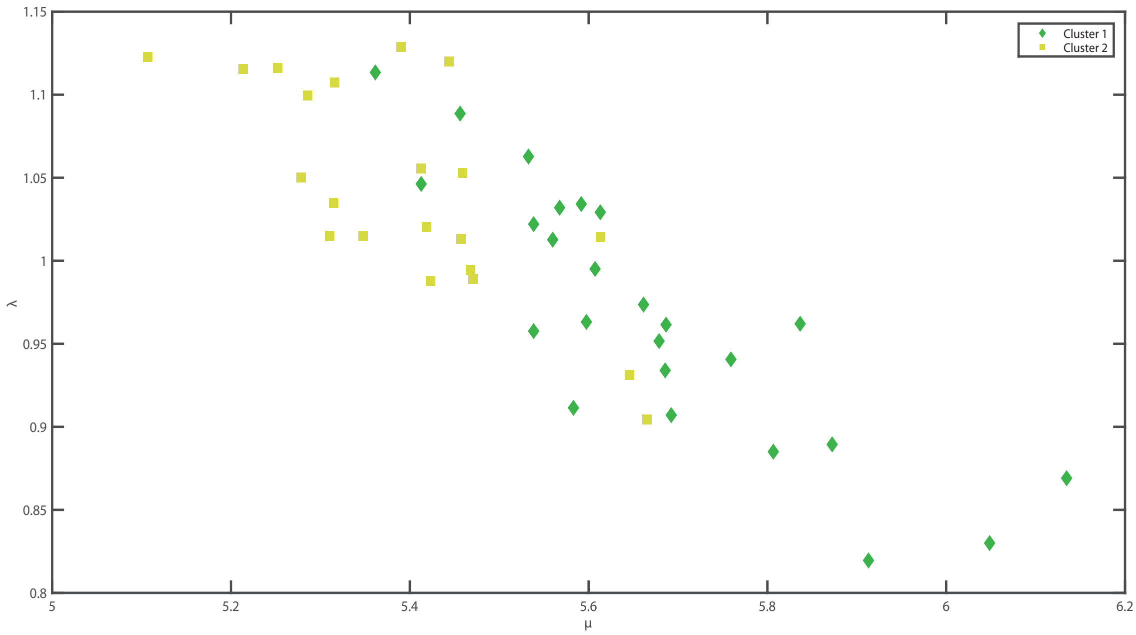

Figure 4. Scatter plot of permanent difference (μ_i) vs. weight to a common factor (λ_i) across prefectures

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, let us 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 was explained in the previous subsection, there are negative relation 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.

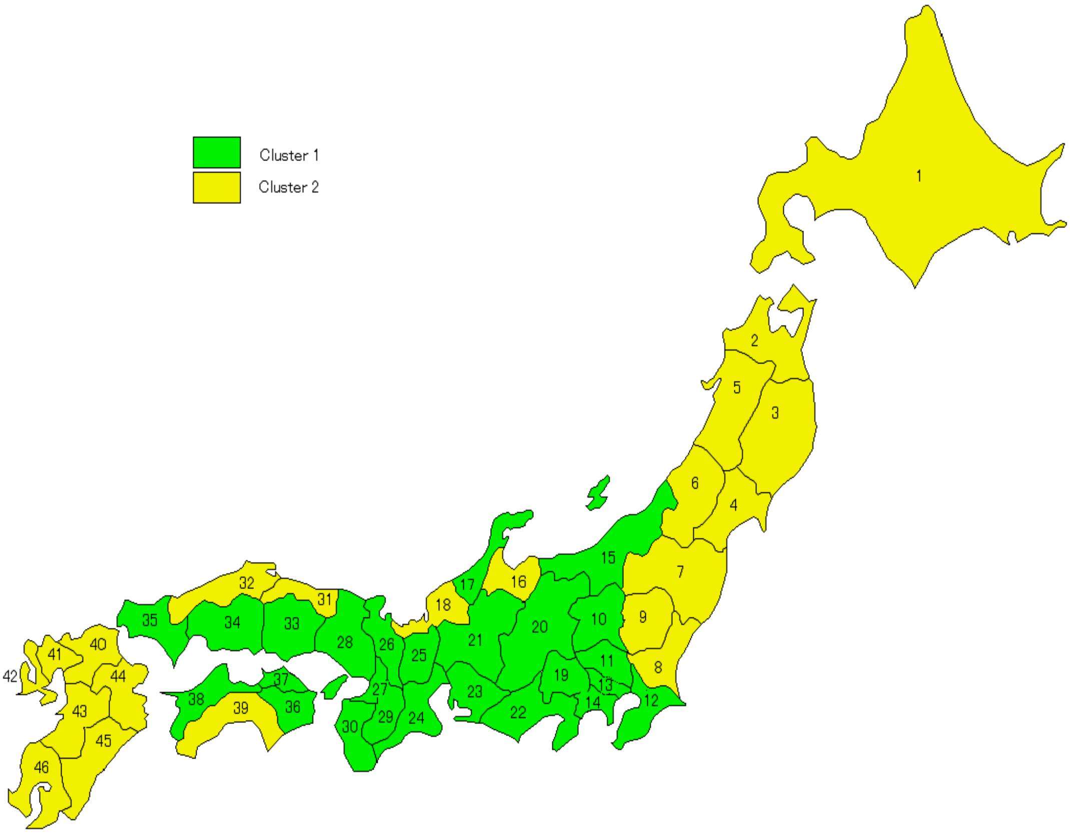

Figure 5. Two clusters of prefectures stratified by the estimates of prefecture-specific components ($\hat{e}_{it}$)

Notes. In this figure, the map of Japan is painted with prefectures of cluster 1 (2) with dark (light) brightness. The clusters are determined 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 represent the prefecture indexes. The correspondence between the index and the prefecture name is given in Table II.

Thirdly, we show that poor prefectures catch up with rich prefectures faster in the recession periods. To see this, 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 also shown with shadows.[22]

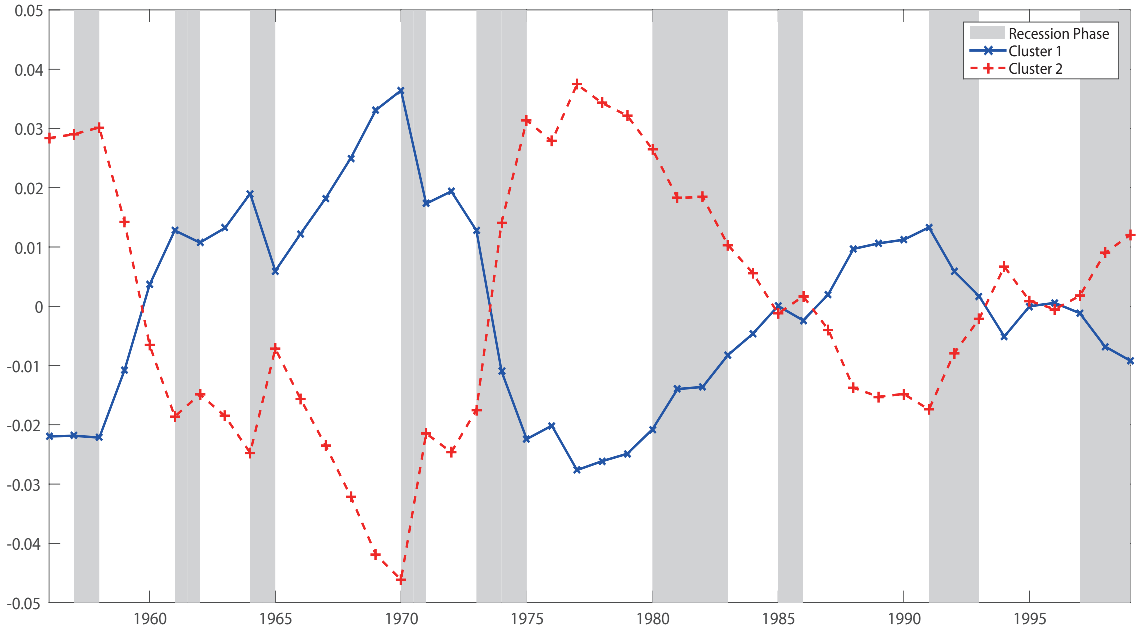

Figure 6. Short-run deviations (e_{it}) of clusters 1 and 2

Notes. The graphs of clusters 1 and 2 are symmetrical by construction, since the total deviation from the average is zero. Shadows in the Figure indicate recession periods.

The gap between the graphs, i.e., e_{it} of cluster 1 − that of cluster 2, represents the short-run income gap between leader-prefectures and follower-prefectures.[23] 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 for the period 1980–1983.

Fourthly, the inequality of regional income should reflect, 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

[22]The data of peaks and troughs are taken from ESRI.
[23]The graphs are symmetrical by construction, since the total deviation from the average is zero.

case, we calculate the Gini index of the logarithm of prefectural per capita income for each year from 1955 to 1999, which is shown in 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 in the short-run gap. This suggests that the income inequality among prefectures *temporarily* widens (shrinks) in an economic boom (recession).

Finally, let us examine whether our results of short-run dynamics are consistent with change in speed of convergence.[24] 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 (high, respectively) in the period of the gap between clusters 1 and 2 increases (decreases).

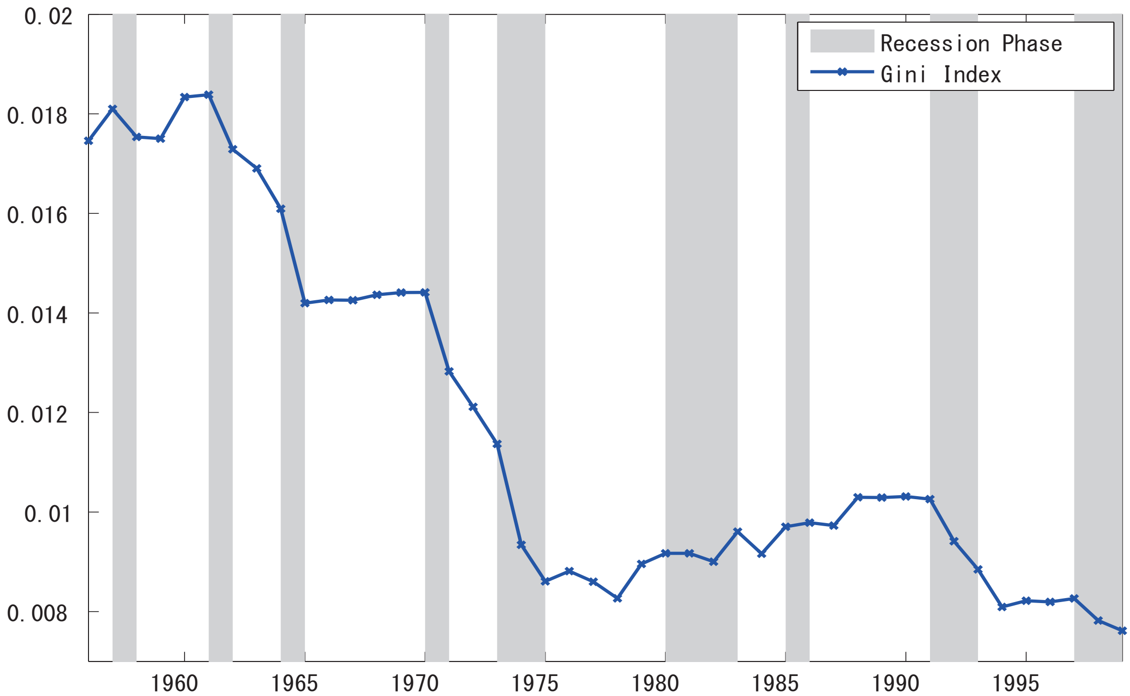

Figure 7. Gini index of the prefectural per capita income

Notes. Shadows in the Figure indicate recession periods.

In order to examine this hypothesis, we need to measure the convergence speed for divided

[24]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.

periods. For the measurement, we split the sample into five-year periods and conduct cross-sectional regressions. Specifically, we estimate the following regression model by nonlinear least squares for each subperiod:

$$\frac{(y_{iT} - y_{i0})}{T} = \alpha - \frac{1 - \exp(-\beta T)}{T} y_{i0} + \epsilon_i, \tag{10}$$

where 0 denotes the initial year and T, the final year for each subperiod. The coefficient β measures the speed of shrinkage of the gap between the initial and final years.

Table V. Nonlinear regressions (10) 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)	

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

Table V presents the estimate of β for 1955–1999 and for the nine subperiods of five years each.[25] According to Table V, 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, in Figure 8, which is the scatter plots of the estimates of β and the changes in the income gap for the nine subperiods, we see a clear negative relationship between them. This suggests

[25]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%).

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.

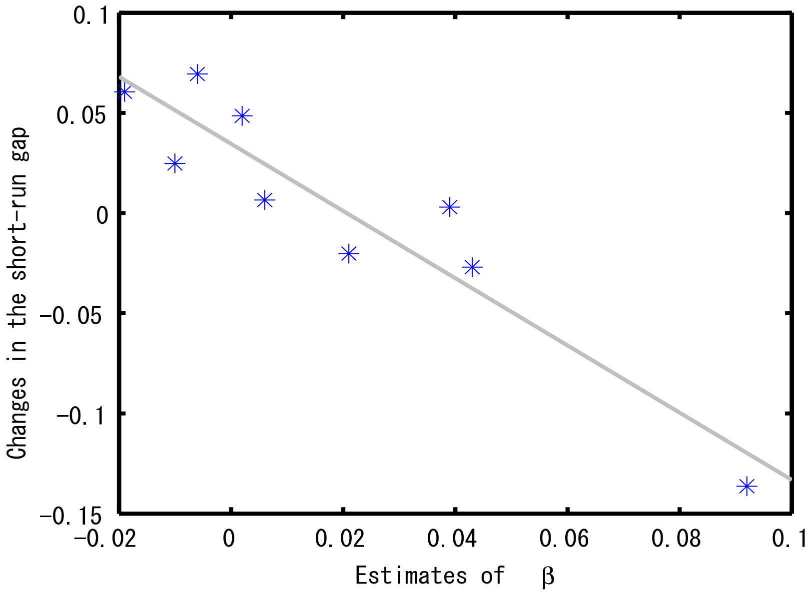

Figure 8. Scatter plot of the convergence speed vs. the changes in the short-run gap between clusters 1 and 2 for nine subperiods

Notes. Nine subperiods are five years each starting from 1955. The fitted line to the nine observations is shown in the figure.

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

5. 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). This method well described the long- and short-run behaviors of regional growth.

Our results are summarized in four 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 one 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 one common factor, F_t, which is almost identical to the mean per capita income of Japan. Then, using unit root and panel cointegration tests, we discovered that the common component contains the stochastic trend characterized by the country-wide labor productivity or TFP and Real Capital, 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 cum smaller coefficient on the common trend than the follower-prefectures, implying that followers will catch up to leaders.

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. This suggests that the income inequality among prefectures *temporarily* widens (shrinks) in an economic boom (recession).

A number of empirical studies have looked at income inequality among Japan's prefectures over a long-run perspective. Most of them argue that the income gap among Japanese prefectures has been decreasing (e.g., Barro and Sala-i-Martin (1992b), Shioji (2001)). Some studies, meanwhile, argue that regional inequality widened in Japan after 2000 (e.g., Tachibanaki and Urakawa (2012), Yamane and Tsutsui (2009)). Our empirical findings can reconcile these stylized and highly controversial facts of real income gap among Japanese prefectures by suggesting the coexistence of both the catch-up process between Japanese leader and follower prefectures in the long-run and the temporal deviation from the long-run equilibrium path. As such, our analysis by the panel cointegration approach in this paper, provides a more realistic picture of the regional income inequality in Japan.

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, how these characteristics for each prefecture come from is still an open question. This constitutes an interesting future task.

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