Yale University

EliScholar – A Digital Platform for Scholarly Publishing at Yale

Discussion Papers

Economic Growth Center

12-1-1991

Regional Growth and Migration: A Japan-U.S. Comparison

Robert J. Barro

Xavier Sala-i- Martin

Follow this and additional works at: https://elischolar.library.yale.edu/egcenter-discussion-paper-series

Recommended Citation

ECONOMIC GROWTH CENTER

YALE UNIVERSITY

Box 1987, Yale Station
New Haven, Connecticut 06520

REGIONAL GROWTH AND MIGRATION:

A JAPAN - U.S. COMPARISON

Robert J. Barro
Harvard University

Xavier Sala-i-Martin
Yale University

December 1991

Abstract

Do poor economies grow faster than rich ones? This important economic question (which we call β-convergence) is analyzed in this paper using two regional data sets: 47 Prefectures in Japan and 48 States of the U.S. We find clear evidence of convergence in both countries: poor prefectures and states grow faster. We also find that there is intra-regional as well as interregional convergence.

We analyze the cross sectional standard deviation across prefectures and states. We find that in both countries there has been a long term decline (a phenomenon that we call σ-convergence).

Finally we study the determinants of the rates of regional in-migration and, again, find striking similarities. In both countries, the reaction of net in-migration rates to the log of initial income is slightly above .025, which indicates a slow (although very significant) speed of population adjustment to income differentials. We find little evidence in favor of the argument that population movements are the reason why we find convergence across economies.

We are interested in the question of whether poor countries or regions tend to grow faster than rich ones. We use the neoclassical growth model as a framework to compare convergence features across 47 Japanese prefectures and 48 U.S. states. The Japanese prefectures and the U.S. states provide clear evidence of convergence in the sense of poor economies tending to grow faster than rich ones in per capita terms. The estimated speed of convergence is consistent with the neoclassical growth model but only if we take a broad view of capital so that diminishing returns to capital set in slowly as an economy develops.

We also want to analyze the question of whether migration has played an important role in this process of convergence. We find some determinants of the rate of in-migration for Japan and the United States and we find that the speed at which population movements react to income differentials (holding constant amenities and population density variables) is very slow in both countries. We also find that the process of convergence is not heavily influenced by internal migration.

(1) Convergence in the Neoclassical Growth Model

The closed economy neoclassical growth models of Ramsey (1928), Solow (1956), Cass (1965), and Koopmans (1965) predict that the per capita growth rate tends to be inversely related to the starting level of output or income per person[1]. In particular, if economies are similar with respect to preferences and technology, then poor economies grow faster than rich ones. Thus, there is a force

---

[1]This result assumes that the production elasticity of substitution between capital and labor is sufficiently bounded away from infinity and the elasticity of intertemporal substitution in the utility function is sufficiently close to constant.

that promotes convergence in levels of per capita product and income. Since the model is well known, we provide only a brief sketch.

We assume that the production function in intensive form is Cobb Douglas with parameter a

$$(1) \qquad \hat{y} = A\hat{k}^{\alpha}$$

where $\hat{y}$ and $\hat{k}$ are output and capital per unit of effective labor, Le^{xt}, L is population or labor, and x is the rate of exogenous, labor–augmenting technological progress. In a closed economy, saving equals investment so $\hat{k}$ evolves as

$$(2) \qquad \dot{\hat{k}} = A\hat{k}^{a} - \hat{c} - (\delta + x + n)\hat{k}$$

where $\hat{c} = C/Le^{xt}$, δ is the rate of depreciation, and n is the growth rate of L. The representative, infinite–horizon household maximizes utility,

$$(3) \qquad U = \int_0^{\infty} [(c^{1-\theta}-1)/(1-\theta)] e^{nt}e^{-\rho t}dt$$

where c = C/L, and ρ is the rate of time preference. We assume $\rho > n + (1-\theta)x$ in the following to satisfy the transversality condition.

The first–order condition for maximizing U in equation (3) entails

$$(4) \qquad \dot{c}/c = (1/\theta)\cdot[a\hat{k}^{a-1} - \delta - \rho]$$

In the steady state, the growth rates of the effective quantities, $\hat{y}$, $\hat{k}$, and $\hat{c}$, is zero and the per capita quantities, y, k, and c grow at the rate x. The level of $\hat{k}$ in the steady state satisfies

$$(5) \qquad \alpha\hat{k}^{*-(1-\alpha)} = \delta + \rho + \theta x$$

Barro and Sala–i–Martin [1991b, Ch. 1]) show that if the economy starts with $\hat{k}$ below $\hat{k}^{*}$, the growth rate of capital per worker, $\dot{k}/k$, declines monotonically toward the steady–state value, x. Because we are assuming a Cobb Douglas production function, this property carries over to the growth rate of output per worker, $\dot{y}/y$. Therefore, if two economies have the same parameters of preferences and technology, then the initially poorer economy—with a lower starting value of $\hat{k}$— tends to grow faster in per capita terms.

We can quantify the transitional dynamics by log–linearizing equations (2) and (4) around the steady state. The average growth rate of y over the interval between dates 0 and T is

$$(7) \qquad (1/T)\cdot\log[y(T)/y(0)] = x + \frac{(1-e^{-\beta T})}{T}\cdot\left[\log[\hat{y}^{*}/\hat{y}(0)]\right]$$

where the positive parameter β, which governs the speed of adjustment to the steady state is given by the formula

$$(8) \qquad 2\beta = \left\{\psi^2 + 4(\frac{1-\alpha}{\theta})(\rho+\delta+\theta x)[\frac{\rho+\delta+\theta x}{\alpha} - (n+\delta+x)]\right\}^{1/2} - \psi$$

where $\psi = \rho - n - (1-\theta)x > 0$.

The higher β the greater the response of the average growth rate to the difference between $\log(\hat{y}^*)$ and $\log[\hat{y}(0)]$; that is, the more rapid convergence toward the steady state. The model implies *conditional convergence* in that for given x and $\hat{y}^*$, the growth rate is higher the lower y(0). Convergence is conditional in that what matters is $\hat{y}(0)$ relative to the steady state values of $\hat{y}^*$ and x, which may differ across economies. In other words, the growth rate of an economy is a decreasing function of the distance between its initial conditions and its own steady state. In cross–country regressions, where we expect economies to be heterogeneous in the sense of having large differences in $\hat{y}^*$ and x, it is important to hold fixed these variations in order to estimate β. One advantage of the regional data we use in this paper is that the differences in $\hat{y}^*$ and x are likely to be minor, so the distinction between conditional and absolute convergence becomes less important.

The complicated expression for β can be greatly simplified if, following Solow (1956), we assume constant gross savings rates. This amounts to restricting the parameters of the model: as shown by Kurz (1968), if the production function is Cobb Douglas and the utility function has a constant intertemporal elasticity of substitution, then value of θ that yields a constant savings rate is[3]

$$(9) \qquad \theta^* = (\delta + \rho)/[a(\delta + n) - x(1-a)]$$

---

[3]See Barro and Sala–i–Martin (1991, chapter 1) for a derivation of this formula.

and the corresponding constant saving rate is

$$(9)' \qquad\qquad s = 1/\theta^*$$

This result is interesting because with a constant saving rate the accumulation constraint can be written as

$$(10) \qquad\qquad \dot{\hat{k}}/\hat{k} = s\hat{k}^{\alpha-1} - (\delta+x+n)$$

We can linearize the growth rate of output around the steady state and get equation (7). The difference is that β is now given by the simple expression

$$(11) \qquad\qquad \beta = (1-\alpha)(\delta+n+x)$$

that is, the speed of convergence is an increasing function of the depreciation, population growth and exogenous productivity growth rates and a decreasing function of the capital share α. In particular, if $\alpha=1$, then the speed of convergence is zero. The case of $\alpha=1$ corresponds to the simplest endogenous growth model: the Ak technology. Note that the speed of convergence in equation (11) is independent of utility parameters (and therefore independent of the saving rate). The reason is that there are two offsetting effects that, for our functional forms, exactly cancel out: given k, a higher saving rate implies a higher speed of convergence. But a higher saving rate also implies a higher steady-state capital intensity, and therefore, a lower marginal product of capital as we approach the steady state. This second effect reduces the speed of convergence.

To assess the quantitative relation between the capital share (which in the Cobb Douglas framework we are using reflects the extent to which diminishing returns set in) and the speed of convergence we use a set of baseline values for the other parameters: ρ=0.05 per year, δ=0.05 per year, n=0.02 per year, x=0.02 per year. The value n=0.02 per year is the average of population growth for the United States over the long history. The other baseline parameters come from estimates reported in Jorgenson and Yun (1986, 1990).

If we adjust the coefficient of intertemporal substitution θ so as to get a constant saving rate along the transition (equation 9), then with a capital share α=0.35 —this is appropriate to a narrow concept of physical capital (see, for example, Maddison [1987])— equation (11) implies β=0.059 per year, which corresponds to a half–life for the log of output per effective worker of 11.8 years. That is, a strict version of the neoclassical model predicts very rapid convergence toward the steady state. The speeds of adjustment that we estimate empirically are much slower: β is in the neighborhood of 0.02 per year. The theory conforms to the empirical findings only if we assume parameter values that depart substantially from the baseline case. For α=0.80, which might apply if capital is interpreted broadly to include human capital, the value β=0.018 per year implies a half–life of 38 years. Of course, as α approaches unity, diminishing returns to capital disappear, β tends to zero, and the half–life tends to infinity.

Open Economy Considerations.

It can be persuasively argued that Prefectures in Japan and States in the United States are *not* closed economies in that inputs of production can move across regions: people can migrate, physical capital can be easily transported, and financial markets allow for borrowing and lending across

regions. When one introduces perfect factor mobility in the neoclassical model just described, one tends to get unrealistic implications such as the prediction that the most patient region owns everything asymptotically and all the other regions have negligible consumption per effective worker. For our purposes, the main implication of perfect capital mobility is the prediction of infinite speeds of convergence.

Barro, Mankiw and Sala-i-Martin (1992) introduce partial capital mobility in a two-capital-good neoclassical growth model. One of the capitals can be used as collateral in international/interregional borrowing and lending and the other cannot (alternatively, one of the capital goods is mobile across regions and the other is not). They show that this open economy version of the neoclassical model can predict speeds of convergence close to .02 for very reasonable parameters. For instance, if the broad capital share is around .8 and if we call λ the fraction of this aggregate measure of capital that can be used as collateral in international credit transactions, the model predicts speeds of convergence between .014 and .035 for a range of λ between 0 and .75: that is, if the amount of capital that can be used as collateral is between 0 and 75%, then the predicted speed of convergence is close to 2% per year. Thus, the fact that the regions we are studying are *not* closed economies should *not* affect our analysis substantially.

Convergence and Endogenous Growth

Finally, even though we put our empirical findings within the framework of the neoclassical model, this is not the only model that can be used. Mulligan and Sala-i-Martin (1991) have recently shown that some two-sector, two-capital goods models of endogenous growth predict conditional convergence

in the same way the neoclassical model does. They show that the conditional convergence coefficients predicted by such models may be similar for very reasonable parameters. Therefore, most of the empirical findings of this paper can also be explained by some two-sector endogenous growth models and should not be taken as evidence against endogenous growth models

(2) Empirical Implementation

Consider a version of equation (7) that applies for discrete periods to economy i and is augmented to include a random disturbance:

$$(12) \quad (1/T)\log(y_{it}/y_{i,t-T}) = x_i^* - \log(\hat{y}_i^*/\hat{y}_{i,t-T})(1-e^{-\beta T})/T + u_{it}$$

where x_i^* is the steady state per capita growth rate and u_{it} is an error term. Although the coefficient β can vary across economies, we neglect these differences in our analysis.[4] This assumption is tenable for the U.S. states, which are likely to be similar in terms of the underlying parameters of technology and preferences. Also, as mentioned before, the theory implies that pure differences in the level of technology do not affect β. Thus, β can be similar for economies that are very different in other respects.

Two Concepts of Convergence

In our earlier papers we discussed two concepts of convergence. The first one, which we called β-convergence, asks the question of whether poor

---

[4]Mulligan (1991) shows that the speed of convergence decreases with $\hat{y}$. We abstract from this problem at the moment.

economies tend to grow faster than rich ones. This is the concept we discussed above. The second one, called σ-convergence, relates to the decline of the cross-sectional dispersion of per capita income or product.

What is the better concept of convergence?. We think that they are both interesting, though different. If we are interested in how fast and to what extent the per capita income of a particular region is likely to catch up to the average across economies, then β-convergence is the relevant concept. If we are interested in the distribution of per capita income across regions, then σ-convergence is what we care about.

We should note that β-convergence is a necessary condition for σ-convergence, but it is not sufficient. That is, a negative coefficient on initial income (a positive β in our notation) does not necessarily imply that the cross-sectional dispersion of per capita output decreases over time. A positive β tends to reduce the dispersion in $\log(y_{it})$ from equation 11, but new shocks, u_{it}, tend to raise it. Equation (11) implies, for a given distribution of u_{it}, that the cross-sectional standard deviation of $\log(y_{it})$, denoted by σ_t, approaches a constant σ. The dispersion falls (or raises) over time if it starts above (or below) the steady state value σ. Hence, β-convergence (positive values of β) need not imply σ-convergence (declining σ_t over time)[5].

In the real world, there are aggregate disturbances, such as wars, and large shocks to agriculture or mineral prices, that affect economies in different ways. For instance, an increase in the price of oil at time t will reduce output (at least in value terms) for the regions that use oil more intensively and increase output for regions that produce oil or use it less

---

[5]See Barro and Sala-i-Martin (1992).

intensively. These disturbances tend to increase the values of σ_t above the steady state σ. After the shock (and assuming that the steady state distribution of u_{it} did not change) there should be a fall in the value of σ_t over time.

If we do not control for these aggregate shocks, then we may obtain biased estimates of β. For instance, imagine an increase in the price of oil at a time in which the oil producing regions are richer (as the U.S. states during the second oil shock). Because of the positive correlation between the aggregate shock and initial per capita income, the estimated β coefficient is biased downward (it will appear that the rich regions grow faster so there no β–convergence). In our econometric analysis we try to hold constant these effects by introducing regional dummies and sectoral variables.

(3) Personal Income Across Japanese Prefectures.

Analysis of β-convergence Across Prefectures and Districts.

We start by analyzing the pattern of β convergence for per capita income across Prefectures[6]. In Table 1 we report regression estimates of the convergence coefficient, β for the period 1930–1987. The first set of columns (headed 'Basic Equation') is an estimation of equation 1 where only the

---

[6]The data on Income are from the Economic Planning Agency (EPA), "Annual Report on Prefectural Income," various issues. The source for the 1930 prefectural income is the "National Economy Studies Association" (literal translation from Japanese). The population estimates are prepared by the Statistical Bureau at the Management and Coordination Agency (MCA). The unavailability of reliable price levels at the Prefectural level forces us to deflate income by country–wide price levels (also prepared by the MCA). Sala–i–Martin (1990) shows that this is not a big problem for the United States. Shioji (1991) uses regional CPI data and argues that the process of converegence is similar.

logarithm of initial income is included. The first row corresponds to a single regression for the whole period 1930–1987. The estimated β coefficient is .0279 (s.e.=.0031) with an adjusted R^2 of .92.   (recall that a positive coefficient corresponds to poor regions growing faster than rich ones). The standard error of the regression is .0020. The amazing fit can also be appreciated in Figure 1. The evidently strong negative correlation between the growth rate 1930–1987 and the log of income per capita in 1930 confirms the existence of β–convergence across the Japanese prefectures.

We can ask the question of whether this convergence process is due to regions catching up or convergence within regions. In Figure 2 we plot the growth rate between 1930 and 1987 for seven japanese Regions or Districts on the log of income per capita in 1930. The definitions of Districts, which are taken from the E.P.A. are displayed in Appendix A. We observe that there has been substantial regional catching up. The point estimate of β for this seven data point regression is .0261 (s.e.=.0079, R^2=92). Thus, the speed of convergence across regions seems to be similar to that across prefectures.

The next few rows of Table 1 break the sample period into six subperiods starting in 1955. We find that for three of the subperiods, the sign is the opposite to the one expected and in one case significant. The speed of convergence is large and very significant for the periods 1960–65, 1970–75 and 1975–80. A constrained estimate for the six subperiods is .0103 (s.e.=.0037). A test for the equality of coefficients over time is strongly rejected (LR statistic=90.23, p–value=0.000). Figure 3 plots the relation between the growth rate between 1955 and 1987 versus the log of income in 1955.

The second set of columns in Table 1 introduces regional dummies for the whole sample and for each subperiod. As mentioned above, these regional dummies proxy for differences in steady state values of per capita income and

also absorb fixed regional effects in the error term, u_{it}. The point estimate for the whole sample is .028 (s.e.=.0024), not significantly different from the one we got when regional dummies were excluded (β=.0279, s.e.=.0031) and not very different from the estimate we got when we used regions only (β=.026, s.e.=.0079). Thus, we confirm that part of the story is convergence across regions and part is convergence across prefectures within regions. When we divide the sample into six subperiods we find now that only the first coefficient has the wrong sign, although it is not significant. The restricted estimate is .0242 (s.e.=.0037). We still reject the equality of coefficients at the 5% level (LR statistic=33.89, p–value=.000).

Additional Variables.

One reason for the apparent instability of the convergence coefficients could be the existence of aggregate shocks. Following Barro and Sala–i–Martin (1991, 1992), the third column of Table 1 adds an additional variable to the regression in an attempt to hold these aggregate shocks constant. We call it S_{it} (for Structure) and it is calculated as follows:

$$() \qquad S_{it} = \sum_{j=1}^{9} \omega_{ij,t-T} \, \log(y_{jt}/y_{j,t-T})/T$$

where $\omega_{ij,t-T}$ is the weight of sector j in state i's personal income at time t–T, and y_{jt} is the national average of personal income per worker in sector j at time t. The sectors used are: agriculture, mining, construction, manufacturing, trade, finance and real estate, transportation, services and government. This structural variable indicates how much a state would grow if

each of its sectors grew at the national average rate. We think of S_{it} as a proxy for common effects related to sectoral composition in the error term, u_{it}. Note that it depends on the contemporaneous growth rates of national averages and on lagged values of own sectoral shares. Suppose that region i specializes in the production of cars and imagine that the aggregate car sector does not grow over the period between t-T and t. The variable S_{it} would then be very low for this region, indicating that, due to some shock that affects the automobile industry, this region should not grow very fast. Hence, we think of this variable as an attempt to hold constant the aggregate shocks we mentioned before.

Due to the lack of data (our prefectural sectoral data start in 1955), we can only include the structural variable for the periods after 1955. Contrary to the previous two columns, none of the subperiods now has the wrong sign (although 1955-60 and 1980-87 are not significant). The restricted coefficient is .034 (s.e.=.0044). The likelihood-ratio test for the equality of coefficients over time is 16.38 with a p-value of .006. Note that it is the period between 1970 and 1975, with a coefficient of .661 (s.e.=.0118) which seems statistically different from the rest. If the oil shock of 1973 hurt the rich industrial regions relatively more, the 1970-75 period would appear as a high convergence period. Somehow, however, the structural variable does not seem to capture this effect fully. We will attempt to break down the sectoral composition variables more thinly in future drafts of the paper.

Analysis of σ-Convergence Across Prefectures and Districts.

We want now to assess the extent to which there has been σ-convergence across prefectures and regions in Japan. We calculate the unweighted cross-sectional standard deviation for the log of per capita income, σ_t, for the 47 prefectures from 1930 to 1987. Figure 4 shows that the dispersion of personal income increased from .47 in 1930 to .63 in 1940. One explanation of this phenomenon is the explosion of military spending during the period. The average growth rate for District 1 (Hokkaido-Tohoku) and 7 (Kyushu), which are mainly agricultural, was -2.4% and -1.7% per year respectively. On the other hand, the industrial regions of Tokyo, Osaka and Aichi grew at +3.7%, +3.1%, and +1.7% per year respectively.

The cross prefectural dispersion decreased dramatically since 1940: it fell to .29 by 1950, to .25 in 1960, to .23 in 1970 and hit a minimum of .125 in 1978. It has increased slightly since then: σ_t rose to .13 in 1980, .14 in 1985 and .15 in 1987.

One popular explanation of the increase in dispersion for the 1980's is the take-off of the Tokyo region from the rest of Japan. Since Tokyo was relatively richer at the end of the 1970s (average per capita income in real terms for Tokyo region was 2.000 billion yen and the average for the rest of Japan was 1.751 billion yen) and grew faster during the 1980s (2.95% a year versus 2.16% a year) that could explain this apparent divergence. To check this point we calculated the cross sectional standard deviation of the log of per capita income for the seven Japanese Districts, and for the six Districts exclusive of Kanto-Koshin (which includes Tokyo). The results are reported in Figure 5. First note that the regional pattern is very similar to the prefectural pattern presented in Figure 4. Second, note that the exclusion of

the Tokyo region shifts the cross sectional variance down, but it does not change the general behavior of σ_t. Third, the increase in dispersion during the 1980s is steeper if the Tokyo region is included, but it is still increasing if excluded. Thus, even though Tokyo contributed to the general increase in cross sectional dispersion during the 1980s, its take off does not fully explain it.

(3) Personal Income Across the U.S. States.

β-Convergence across U.S. states and regions.

It is interesting to compare the results on convergence for Japanese prefectures with those for the U.S. states (see also Barro and Sala-i-Martin (1991b)). The sample available for the Unites States begins in 1880.

Figure 6 shows the growth rate between 1880 and 1988 versus the logarithm of income in 1880. As it was the case for Japan, the negative correlation (reflecting β convergence) can be captured by the naked eye. The point estimates for convergence are also similar. For the overall sample, the estimated β is .0171 (s.e.=.0028).

Table 2 reproduces Table 1 using personal income per capita across the 48 contiguous states of the United States.[7] The sample period (1880–1988) is divided into nine subperiods. The point estimates for the simple convergence equation (exclusive of regional dummies and structural variables) has the wrong sign for two of the nine subperiods: 1920–1930 (the period of large agricultural price changes) and 1980–1988 (the period following the oils

---

[7]The early data are from Easterlin (1960a, 1960b). Data after 1920 are from Bureau of Economic Analysis, *Survey of Current Business*, various issues.

shocks). The restricted point estimates for this basic equation is
.0175 (s.e.=.0013). We reject the hypothesis of stability of the β
coefficients over time as the LR statistic is 65.6 with a p–value of 0.000.

When we include regional dummies, only the 1920s have the wrong sign.
The restricted point estimate of β is .0189 (s.e.=.0019). The likelihood
ratio statistic for the test of equality of coefficients over time is 32.1 so
we still reject the hypothesis of stability of β over time. The fact that β
is positive after holding constant the regional dummies indicates that there
is substantial within–region convergence. To further investigate the
possibility of between–region convergence, we constructed regional income
variables for the four main U.S. regions. The plot of growth between 1880 and
1988 and the log of per capita income in 1880 is reported in Figure 7. The
negative relation is, again, self evident. Hence, as it was the case in
Japan, there appears to be between– as well as within–region convergence.

When the structural variables[8] are included, the restricted point
estimate for β=.0224 (s.e.=.0022). The likelihood–ratio statistic is 12.4.
The critical value at the 5% level is 15.5 so we cannot reject the hypothesis
of stability of β over time at the 5% level.

σ-*convergence* across *U.S. states and regions.*

Figure 8 shows the cross–sectional standard deviation for the log of per
capita personal income for 48 U.S. states from 1880 to 1988. We observe that
the dispersion declined from .54 in 1880 to .33 in 1920, but then rose to .40
in 1930. This rise reflects the adverse shock to agriculture during the

---

[8]Since data on structural shares were available only after 1929, we
measured S_{it} by the agriculture share in total income for the earlier period.

1920s: the agricultural states were relatively poor in 1920 and suffered a further reduction in income with the fall in agricultural prices. The U.S. value for 1930 is similar to (although slightly below the) value for Japan for the same year (σ_t=.47). After 1930, the U.S. σ_t fell to .35 in 1940, .24 in 1950, .21 in 1960, .17 in 1970, and a low point of .14 in 1976. As it was the case in Japan, the long run decline stopped in the mid–1970s, after the oil shock, and σ_t rose to .15 in 1980 and .19 in 1988. Note that, after 1940, the patterns of the dispersion in Japan and the United States have been highly correlated : there is a long term decline with a reversal somewhere in the middle of the 1970s, and the low point for σ_t is similar (.14 for the U.S. and .12 for Japan). Of course one possible explanation is that the steady state value for σ is close to .12 (or maybe a bit higher). Under this hypothesis, it would not be surprising to see that, after a long period of decline, σ_t remains close to this level for a decade or so.

The main lesson is that the convergence patterns across states and regions of the United States and prefectures and regions of Japan appear to be very similar. There is convergence across as well as within regions and the speed of convergence is similar, although slightly higher in Japan: the restricted coefficients are .0189 (s.e..0019) for the U.S. and .0242 (s.e.=.0037) for Japan when we just hold constant regional dummies and .0224 (s.e.=.0022) for the U.S. and .0340 (s.e.=.0044) for Japan when we also include sectoral variables. In terms of σ–convergence, the similarity between Japan is even more striking: after a long term decline in σ_t, they both hit a minimum somewhere during the mid to late 1970s and they both have been increasing since. The minimum value of σ is very close: .14 for the U.S. and .12 for Japan.

(4) Net Migration across Japanese Prefectures.

It can be argued that the convergence across states, prefectures and regions is only due to population movements. In this section we want to see whether the patterns of convergence found in the last sections can be explain by migration. We start with the study of the determinants of net in-migration in Japan and the United States. We then combine this analysis with that of the previous sections to conclude that migration is unlikely to be the main source of interregional convergence.

In Barro and Sala-i-Martin (1991a) we discussed the modeling of migration within the context of the neoclassical growth model (see also Mueser and Graves [1990]). We arrived at a reduced form expression for m_{it}, the annual rate of net migration into region i between years t-T and t.

(14) $m_{it} = f(y_{i,t-T}, \theta_i, \tau_{i,t-T}$, and variables that depend on t but not i)

where $y_{i,t-T}$ is per capita income at the beginning of the period, θ_i is a vector of underlying amenities (such as climate, geography, etc. which do not change over time), and $\tau_{i,t-T}$ is the population density in region i at the beginning of the period. The set of variables that depend on t but not on i includes any elements that influence the national averages of per capita income and population densities. The set also includes effects like technological progress in heating and air conditioning and in transportation from the suburbs to the center of the city, which affect people's attitudes towards weather and population densities.

We expect initial income to have a positive (partial) effect on migration and densities to affect migration negatively. When trying to implement this

relation for Japan, we should realize that there is a substantial difference between a Japanese prefecture and a U.S. state. The average size of a Japanese prefecture is 6,394 square Kilometers[9], roughly half the size of Connecticut. The largest prefecture, Hokkaido, is 83,520 Km^2 or roughly the size of South Carolina. The second largest prefecture, Iwate, has an area of 15,277 Km^2, a bit larger than Connecticut and a bit smaller than New Jersey. In comparison, the average U.S. State has an area of 163,031 Km^2 and the area of the largest state in the continental U.S., Texas, is 691,030 Km^2. California, with an area of 411,049 Km^2 is slightly larger than all Japan (377,682 Km^2).

This of course means that, unlike states of the United States, Japanese prefectures look more like cities and, therefore dealing with day-time commuting is important. Researchers in the urban economics literature[10] think that people like cities for two reasons. First there are demand or consumption externalities. That is, cities provide amenities such as theaters, opera houses, museums, etc. that can be supplied only if there is large demand for them. Second, there are production externalities that tend to increase wages around big cities. On the other hand, people do not like crowded cities because they tend to be associated with crime, less friendly neighborhoods, low quality of living and, more importantly, high land and housing prices (see Roback [1982]). Thus, when deciding when and where to

---

[9]Excluding Hokkaido, which is at least eight times as large as all the others. The average size including Hokkaido is 8,036 Km^2, two thirds the size of Connecticut.

[10]See for instance Henderson (1988).

migrate, people face a trade off. If people are allowed to move to an area close to a crowded city and commute from there, they will be able to enjoy the benefits without paying the costs (equilibrium is reached because people have to pay a large commuting cost in exchange for the benefits). People will be especially willing to pay commuting costs when densities are extremely large.

Ideally, to deal with these issues empirically we would like to have a measure of "density of the neighboring prefectures". Conceptually, we could construct such a measure by weighting the neighbor's density by their distance in some (possibly exponential) way. In practice, however, we observe that there are two areas with an abnormally high population density: Tokyo and Osaka. In 1987, Tokyo's density is 5,400 people/Km^2 and Osaka's is 4,567 people/Km^2. The average for the rest of the prefectures in the same year is 415 people/Km^2.[11] Hence, the problems mentioned above are likely to arise in these two regions only. We can confirm this statement by looking at actual commuting numbers":[12] we compute the ratio of day-time to night-time population. A ratio smaller than one indicates that there are people who live in that prefecture but work in another while a ratio larger than one indicates the opposite. We find that the ratio is basically one for all prefectures except for the ones around the Tokyo and Osaka areas: Tokyo's ratio is 1.184 and Osaka's is 1.053. The ratios for the Tokyo region are .872 for Saitama, 87.6 for Chiba and .91 for Kanagawa. For the Osaka region the ratios are .955 for Hyogo, .871 for Nara and .986 for Wakayama.[13] We constructed a variable

---

[11]In comparison, the U.S. state with the largest density in 1987 was New Jersey with 273 people/Km^2.

[12]Source: Statistics Bureau, Management and Coordination Agency.

[13]There seems to be some commuting in Kyoto and Aichi, although the magnitudes are certainly much smaller: Aichi's ratio is 1.016 (and its neighboring prefecture, Gifu, has a ratio of .977) and Kyoto's is 1.011.

called "Neighbor's Density" by assigning to every prefecture its own density and the prefectures of the Tokyo area (Tokyo and its immediate neighbors Saitama, Chiba and Kanagawa) and Osaka area (Osaka and its immediate neighbors Hyogo, Nara, and Wakayama) the density of their area[14]. We expect to get a positive partial relation between migration and this neighbor's variable and a negative partial relation between migration and own density. This relation would indicate that people do not like to live in dense areas (they have to pay congestion costs) but they like to be close to them (so they can still get the benefits of a big city).

The empirical function we implement is the following:

$$(15) \quad m_{it} = a + b \cdot \log(y_{i,t-T}) + c_1 \theta_i + c_2 \tau_{i,t-T} + c_3 \tau^{NE}_{i,t-T} + v_{it}$$

where v_{it} is an error term, and $\tau^{NE}_{i,t-T}$ is the population density of the surrounding prefectures. In order to implement this relation empirically, we calculated population densities by dividing total population (in thousands of people) by total area (in square kilometers). To calculate the amenity (weather) variable we squared the difference between the maximum and the average temperatures and we added the square of the difference between the minimum and the average. We took the square root of the result. Hence, this variable measures extreme temperature. A variable similar to the one used for the United States (heating degree-days) was unavailable. We experimented with other weather variables such as maximum, minimum temperatures, or average snow fall over the year. These alternative variables did not fit as well.

---

[14]Because the densities for the rest of the prefectures are very similar, we think that this approximation is good enough (ie, the average of the neighbors and the own density is almost the same).

Figure 10 shows the relation between the average annual migration rate 1955–1987 and the log of income per capita in 1955. The clear positive association (the simple correlation is .58) suggests that net migration reacts positively to income differentials. An interesting point is that the three outliers at the top of the figure are Chiba, Saitama and Kanagawa: the three prefectures surrounding Tokyo.

Table 3 shows the results of estimating migration equations of the form in equation (15). The first row refers to the average migration rate for the whole period 1955–1987. The coefficient for the log of initial income is .0173 (s.e.=.0044). As expected, own density is negatively associated with net migration (−.0042, s.e.=.0017) and neighbor's density is positively associated (+.0092, s.e.=.0016). The extreme temperature variable is negative although its level of significance is far from impressive.

The next six rows in Table 3 show results for the six subperiods beginning with 1955–60. The coefficient on initial income is significantly positive for all subperiods (the exception is 1975–1980, which is positive although not significant). The restricted estimate is .0271 (s.e.=.0025). which implies that, other things equal, a 10 percent increase in a prefecture's per capita income raises net in-migration (only) by enough to raise that prefecture's rate of population growth by .27 percentage points per year. This is a slow speed of adjustment.

Again, as expected, the own density variable is significantly negative and the neighbors' density variable is significantly positive (the exception is the initial period). The constrained estimates are −.0039 (s.e.=.0005) and .0058 (.0008) respectively. The extreme weather variable is negative but not very significant. The constrained estimate is −.0004 (s.e.=.0002). Thus, weather does not seem to play an important role in the process of internal

migration in Japan.

The main findings of this section are that the rate of net in-migration to a prefecture is strongly negatively related to the own density and strongly positively associated with the density of the neighbors. Holding other things constant, migration is strongly positively associated with initial income. The point estimate for the speed of migration is around .025: the constrained estimate is .027 (s.e.=.0025), which implies that, other things equal, a 10 percent increase in a prefecture's per capita income raises net in-migration (only) by enough to raise that prefecture's rate of population growth by .27 percentage points per year. Hence, even though the coefficient is strongly significant, it is quite small in magnitude. This slow adjustment means that population densities do not adjust rapidly to differences in per capita income. Our previous results suggest that differences in per capita income tend to vanish at a slow speed themselves (between 2.5% and 3% per year). Putting these two results together, the implication is that net migration rates are highly persistent over time. The data confirm this conclusion: the correlation between the average migration rate for the period 1955–1970 and that for the period 1970–1987 is .60.[15]

We can gain further grasp on the patterns of internal migration by looking at the time series rates of in-migration for each of the seven Districts for the period 1955–1987. Figure 11a shows that Districts 1, 5, 6 and 7 (Hokkaido–Tohoku, Chugoku, Shikoku, and Kyushu) experienced negative net migration throughout the period. Out-migration from these districts was substantial between 1955 and 1970 (losses of about 1% of the population every

---

[15]The correlation excluding Tokyo and Osaka is even higher, .77. The reason is that Tokyo and Osaka experience large positive migration rates in the first period and, due to the congestion effects described above, they experience large negative migration rates in the second half of the sample.

year), but it has slowed down since. In terms of absolute population figures, Hokkaido-Tohoku and Kyushu were losing about 150,000 people per year, while the losses for Chugoku and Shikoku were closer to 50,000 people per year. Out-migration after 1970 has been much smaller (between -.5% and 0% per year).

Figure 11b shows that most of the positive net migration rates during the 1950s and 1960s occurred in the Districts of Kanto-Koshin (number 2, which includes Tokyo) and Kinki (number 4, which includes Osaka and Kyoto). The net migration figures, however, dropped dramatically in the beginning of the 1970s, coinciding with the first increase of land prices. Kinki has experience negative but small net migration rates since then. Kanto-Koshin's migration rates remained positive during the 1970s and 1980s and have been increasing again since 1976.

The District of Kanto-Koshin includes Tokyo as well as its neighbors. As we saw in the regression analysis above, the behavior of these two sets of prefectures is likely to be very different. In Figure 11c we compare the pattern of migration for Tokyo with that of its three immediate neighbors (Chiba, Saitama, and Kanagawa). Note that net migration rates in Tokyo were positive and very large in the 1950s (the contribution of migration to population growth in Tokyo was between 2% and 3% per year). These enormous rates were falling rapidly and became negative in the second half of the 1960s. It has stayed negative ever since, even though the fall in net migration rates stopped somewhere in the middle of the 1970s.

The pattern for the three neighbors has been strikingly different. After a near zero net migration rate in 1955, migration picked up strongly and reached a maximum of 3% per year in 1963. It remained around 3% until 1972, at which time they started a slow decline. They seem to have stabilized around 1% per year in the 1980s. Thus, there seems to be a pattern of

out-migration from Tokyo into its immediate neighbors.

(5) Net Migration Across the U.S. States.

Again, it is interesting to compare these migration patterns with those for the states of the U.S. As mentioned above, since the size of the average prefecture is a lot smaller than that of the average U.S. state, the congestion/commuting problems that arise in Japan are less important for the U.S. state data. Our U.S. state net migration data start in 1900 and are available every census year (except for 1910 and 1930). We calculate the ten year annual migration rates by dividing net migration between t-T and t by the stock of population at t-T.

Figure 12 shows the simple, long term relation between migration and the log initial income per capita[16]. The horizontal axes plots the log of the state per capita income in 1900. The positive association is evident (correlation=.51). The main outlier is Florida, with a lower than average initial income per capita and a very high net migration rate of 3% per year.

Table 4 shows regression results similar to those for Japan. The dependent variable is the net rate of in-migration m_{it}. The functional form we found to fit best for the United States (see Barro and Sala-i-Martin (1991a)) includes the period specific coefficients for $\log(y_{i,t-T})$, single coefficients for density and the square of density, and period specific coefficients for the log of heating degree-days. The regressions also include period specific coefficients for regional dummies and agricultural shares (the

---

[16]The variable on the vertical axis is the average annual in-migration rate for each state from 1900 to 1987. The variable is the average for each subperiod weighted by the length of the interval.

estimated coefficients for the latter two variables is sometimes significant,
but play a very minor role overall).

The estimated coefficients for $\log(\text{Heat}_i)$ in Table 4 are all negative and
most are significantly different from zero: other things equal, people like
warmer states. The jointly estimated linear term for density is –.0452
(s.e.=.0077) and the square term is significantly positive, .0340
(s.e.=.0092). These point estimates imply that the marginal effect of
population density on migration is negative for all states except for the
three with highest densities (New Jersey, Rhode Island since 1960, and
Massachusetts since 1970).

The coefficient on the log of initial income is strongly significantly
positive for all subperiods. The joint estimate is .0261 (s.e.=.0023), which
implies a t–value over 11. Hence, there is strong evidence that, other things
equal, migration relates positively to initial income. The estimates of Table
4, however, reject at the 5 percent level the hypothesis of stability of the
income coefficients over time.

Note that the main results for Japan and the United States are similar:
people move away from highly populated areas and into high income areas.
Thus, to the extent that there are scale benefits from population densities,
these have to show up as higher per capita income to get people to @@ the
congestion (and higher housing prices). Bad weather is bad for migration
(although this does not seem to be an important factor for Japanese
migration). One striking similarity is the small but significant coefficient
on initial income: .026 for the United States (s.e.=.0023) and .027 for Japan
(s.e.=.0025). These results mean that, other things equal, a 10% differential
in income per capita raises net in–migration only by enough to raise the
area's rate of population growth by .26% per year in the United States and

.27% per year in Japan.

(6) Convergence and Migration.

Migration of low human capital people from poor to rich regions speeds up the convergence of per capita income. If this process has been occurring in Japan and the United States, then the convergence coefficients in Tables 1 and 2 partly reflect the impact of migration. But if migration is a source of convergence, then the introduction of migration rates into the convergence equations would reduce the size of the β coefficient (and if migration is the whole story, then the estimated β would be close to zero once migration is held constant). To quantify the effects of migration on the process of convergence, we enter migration rates into the growth rate regressions of Tables 1 and 2.   For Japan, we find that the joint point estimate for the convergence coefficient is $\beta=.0375$ (s.e.=.0044), which is actually higher than the one we got in Table 1. Moreover, the coefficient on migration is positive and borderline significant (which is the opposite to the expected effect if migration is the cause of regional convergence).

The positive coefficient on the migration variable in the growth equation is likely to reflect the simultaneous determination of migration and per capita growth. If a region has favorable growth prospects that are not fully captured by the explanatory variables that we included in the the regressions for growth and migration, then the residuals in each equation would tend to be positive (the positive residual in the migration equation would reflect the positive growth prospect that are not adequately captured by the other explanatory variables in the growth regression).

We have estimated the growth-rate equation using instrumental variables.

In addition of the predetermined variables already included in Table 1, we include as instruments the variables that explain migration rates in Table 2: the extreme weather variable and the own and neighbors' density variables. If the coefficient on m_{it} in the growth equation is constrained over all subperiods, then the estimated β coefficient is .0363 (s.e.=.0078), not substantially different from the ones we got in Table 1. The migration variable has a positive but insignificant coefficient (.042, s.e.=.189). This result suggests that exogenous shifts in migration do not affect growth significantly: holding constant these exogenous shifts in migration rates, the convergence coefficients are basically the same.

The results for the relation between migration and convergence across the United States are fairly similar. If we include migration as a right hand side variable in the growth regression, the restricted β convergence coefficient, β, is 0.025 (s.e.=.003). This point estimate is (again contrary to what one would have expected and similarly to what we got for Japan) higher than the one we got when the migration rates where excluded. The estimate of the coefficient on migration is 0.098 (s.e.=.029), which is positive and significant.

The instrumental variables estimation for the United States (using the explanatory variables of Table 4 as instruments) suggest that the role of exogenous shifts in migration is not important in explaining convergence across U.S. states: the joint estimate for the instrumented migration coefficient is not significantly different from zero (.010, s.e.=.047) and the convergence coefficient remains unchanged (β=.0214, s.e.=.0030). The findings are the same if we allow for different coefficients for each subperiod.

The main conclusion is that, even though the process of migration is a potential explanation of the convergence process in the United States and

Japan, we do not find evidence of its importance in either case.

(7) Conclusions.

We have compared the process of regional economic growth and convergence across states in the United States and prefectures in Japan. The main conclusion is that the patterns of regional growth are similar for the two countries: There has been a steady reduction of the cross sectional variance of the logarithm of per capita income in Japan and the United States (a phenomena that we call σ-convergence). The process seemed to stop between 1975 and 1980 and the minimum value of the standard deviation is strikingly similar: .13 for Japan and .14 for the United States.

The speed at which poor regions tend to catch up with rich ones is also similar: we estimated the β coefficient to be .025 for the United States and slightly above .030 for Japan. The neoclassical model can be reconciled with these low speeds of convergence if we think of capital in a broad sense so as to incorporate human and other forms of non–physical capital: the required capital share has to be larger than 0.7.

We found that the reaction of migration to regional income differentials in both countries was significantly positive. The size of the coefficient, however, indicated a slow reaction of migration to income differentials: holding constant other determinants of migration, a 10% differential in income per capita raises net in–migration only enough to raise the area's rate of population growth by .26% per year in the United States and .27% per year in Japan. The main difference in the determinants of migration rates between Japan and the United States was the variable "Density of the Neighboring Prefectures," which was significantly positively in Japan. Holding constant

this measure, we found that a prefecture's own density influenced migration in a significantly negative way. This finding indicates that people like to move close to highly populated areas (so that they can enjoy the benefits of high wages and consumption externalities) but, to the extent possible, avoid the costs of highly congested areas (with their crime, high land and housing prices, etc.). If commuting is possible, then people move to low density areas close to high density cities.

The reason for the difference from the United States is the size of a prefecture —about 20 times smaller than the average U.S. state and much closer in size to a city. Issues relating to day-time commuting are important for Japanese prefectures but not for U.S. states.

Finally we showed that exogenous changes in migration seem to be unimportant in explaining the process of interregional convergence.

If we put these results together with those of our previous studies, then we see that the speeds of convergence across Japanese Prefectures and Regions are similar to the speeds of convergence we found for European Regions within the countries of Germany, France, the United Kingdom, the Netherlands, Belgium, , Italy, and Denmark, and the United States, and the conditional convergence that we found for the 20 original OECD countries, and about 100 countries in the Summers and Heston (1991) data set. We are left with a stylized fact that theories of economic growth should explain: a variety of data sets show that economies (conditionally) converge to each other at a speed of about 2% per year.
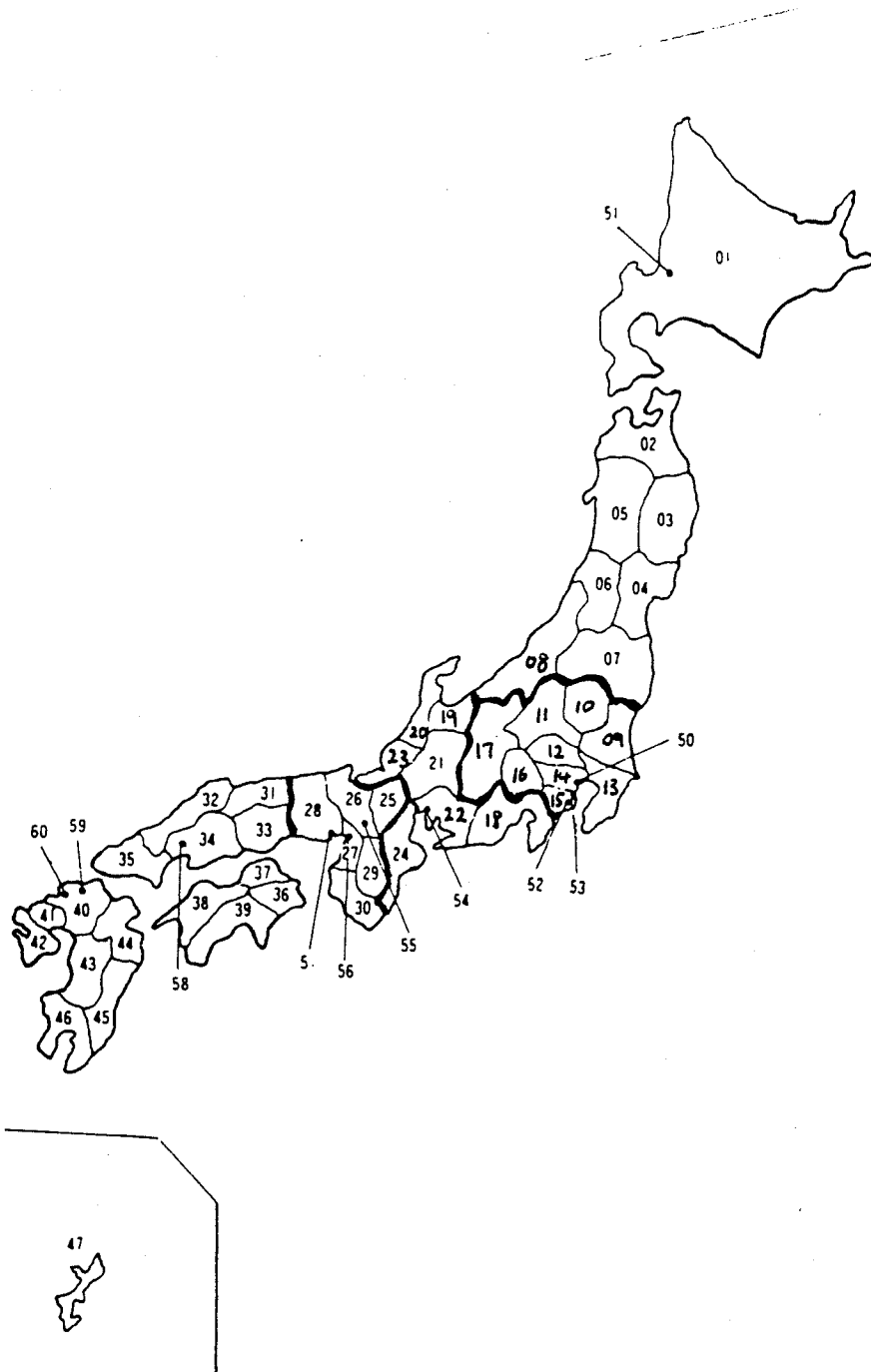
Appendix A:

Prefectures: Key to the Map

1 — Hokkaido	13 — Chiba	25 — Shiga	37 — Kagawa
2 — Aomori	14 — Tokyo	26 — Kyoto	38 — Ehime
3 — Iwate	15 — Kanagawa	27 — Osaka	39 — Kochi
4 — Miyagi	16 — Yamanashi	28 — Hyogo	40 — Fukuoka
5 — Akita	17 — Nagano	29 — Nara	41 — Saga
6 — Yamagata	18 — Shizuoka	30 — Wakayama	42 — Nagasaki
7 — Fukushima	19 — Toyama	31 — Tottori	43 — Kumamoto
8 — Niigata	20 — Ishikawa	32 — Shimane	44 — Oita
9 — Ibaraki	21 — Gifu	33 — Okayama	45 — Miyazaki
10 — Tochigi	22 — Aichi	34 — Hiroshima	46 — Kagoshima
11 — Gumma	23 — Mie	35 — Yamaguchi	47 — Okinawa
12 — Saitama	24 — Fukui	36 — Tokushima	

EPA District Classification

District Number	Name	Prefectures Included
1	Hokkaido–Tohoku	1– 8
2	Kanto–Koshin	9–17
3	Chubu	18–24
4	Kinki	25–30
5	Chugoku	31–35
6	Shikoku	36–39
7	Kyushu	40–47

Appendix B:

States of the U.S.
Abbreviations Used in the Figures

AL – Alabama	IA – Iowa	NE – Shiga	RI – Rhode Island
AZ – Arizona	KS – Kansas	NV – Kyoto	SC – South Carolina
AR – Arkansas	KY – Kentucky	NH – Osaka	SD – South Dakota
CA – California	LA – Louisiana	NJ – New Jersey	TN – Tennessee
CO – Colorado	ME – Maine	NM – New Mexico	TX – Texas
CT – Connecticut	MD – Maryland	NY – New York	UT – Utah
DE – Delaware	MA – Massachusetts	NC – North Carolina	VT – Vermont
FL – Florida	MI – Michigan	ND – North Dakota	VA – Virginia
GA – Georgia	MN – Minnesota	OH – Ohio	WA – Washington
ID – Idaho	MS – Mississippi	OK – Oklahoma	WV – West Virginia
IL – Illinois	MO – Missouri	OR – Oregon	WI – Wisconsin
IN – Indiana	MT – Montana	PA – Pennsylvania	WY – Wyoming

Census Regional Classification

Region Name	States Included
Northeast:	ME, NH, VT, MA, RI, CT, NY, NJ, PA.
South:	DE, MD, VA, WV, NC, SC, GA, FL, KY, TN, AL MS, AR, LA, OK, TX.
Midwest:	MN, IA, MO, ND, SD, NE, KS, OH, IN, IL, MI, WI.
West:	MT, ID, WY, CO, NM, AZ, UT, NV, WA, OR, CA.

Appendix Figure B: The States of the United States.
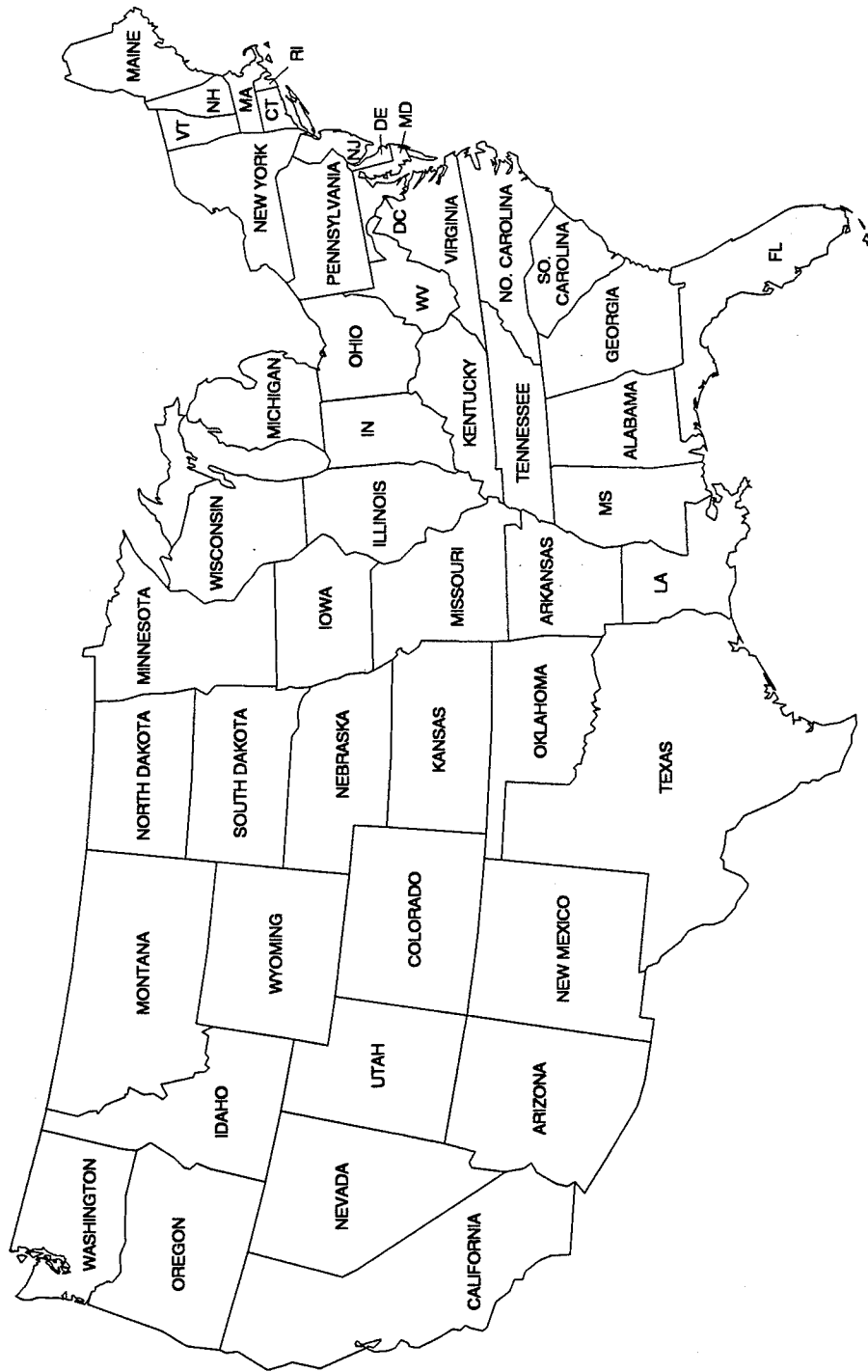
References

Barro, R. J. and X. Sala–i–Martin (1991a), "Convergence Across States and Regions", BPEA, 1:1991, 107–158.

Barro, R. J. and X. Sala–i–Martin (1991b), "Economic Growth", unpublished manuscript.

Barro, R. J. and X. Sala–i–Martin (1992), "Convergence", forthcoming in Journal of Political Economy", 1992.

Barro, R. J., N. Gregory Mankiw, and X. Sala–i–Martin (1991), "Convergence and Capital Mobility", mimeo Harvard University, December 1991.

Baumol, William J. (1986), "Productivity Growth, Convergence, and Welfare: what the Long–Run Data Show", American Economic Review 76: 1072–85.

Blanchard, O. J. and S. Fischer (1989), "Lectures in Macroeconomics", M.I.T. Press, Cambridge MA.

Bureau of the Census, (1975), "Historical Statistics of the United States, Colonial Times to 1970", Bicentennial edition, parts 1 and 2. Washington:

U.S. Department of Commerce, (1990), "Statistical Abstract of the United States: 1990". Washington: U.S. Department of Commerce.

Cass, D. (1965), Optimum Growth in an Aggregative Model of Capital Accumulation", Review of Economic Studies 32: 233–40.

Easterlin, Richard A. (1960a), "Regional Growth of Income: Long Run Tendencies". in Population Redistribution and Economic Growth, United States, 1870–1950. II: Analyses of Economic Change, edited by Simon Kuznets, Ann Ratner Miller, and Richard A. Easterlin. Philadelphia: The American Philosophical Society.

Easterlin, Richard A. (1960b), "Interregional Differences in Per Capita Income, Population, and Total Income, 1840–1950", In Trends in the American Economy in the Nineteenth Century. A report of the National Bureau of Economic Research, New York. Princeton: Princeton University Press.

Economic Planning Agency (Japan), various issues. "Annual Report on Prefectural Accounts".

Henderson, J. Vernon (1988), "Urban Development: Theory, Fact and Illusion", Oxford University Press, Oxford.

Jorgenson, Dale W. and K Yun, (1986), "Tax Policy and Capital Allocation", Scandinavian Journal of Economics, 88, 355–377.

Jorgenson, Dale W. and K Yun, (1990) "Tax Reform and U.S. Economic Growth", Journal of Political Economy, 98.

Koopmans, Tjalling, C. (1965), "On the Concept of Optimal Economic Growth", In Study Week on the Econometric Approach to Development Planning". Pontificiae Academiae Scientiarum Scripta Varia, no 28. Chicago: Rand McNally.

Kurz, M. (1968), "The General Instability of a Class of Competitive Growth Processes", Review of Economic Studies, March, 155.

Maddison, A. (1987), "Growth and Slowdown in Advanced Capitalist Economies: Techniques of Quantitative Assessment", Journal of Economic Literature, 25, June.

Mankiw, N. Gregory, David Romer, and David N. Weil (1991), "A contribution to the Empirics of Economic Growth" forthcoming in Quarterly Journal of Economics", Harvard University, 1992.

Mueser, Peter R. and Philip E. Graves (1990), "Examining the Role of Economic Opportunity and Amenities in Explaining Population Redistribution", working paper 90-4, University of Missouri-Columbia (October).

Mulligan, Casey B. (1991), "A Note on the Time Elimination Method for Solving Nonlinear Dynamic Systems", NBER Working Paper, December 1991.

Mulligan, Casey B. and Xavier Sala-i-Martin (1991), "Transitional Dynamics in Two-Sector Models of Endogenous Growth", mimeo Yale University, December 1991.

Ramsey, Frank P. (1928), "A Mathematical Theory of Saving", Economic Journal, 38, December 1928, 543-559.

Roback, J. (1982), "Wages, Rents, and the Quality of Life", Journal of Political Economy 90: 1257-78.

Sala-i-Martin, Xavier (1990), "On Growth and States" Ph.D. Dissertation, Harvard University.

Shioji, Etsuro (1991), "Regional Growth in Japan", mimeo Yale University, 1991.

Statistics Bureau, Management and Coordination Agency, Japan "Vital Statistics" various issues.

Statistics Bureau, Management and Coordination Agency, Japan, "Statistical Survey on Internal Migration in Japan". Various Issues.

Solow, Robert M. (1956), "A Contribution to the Theory of Economic Growth". Quarterly Journal Of Economics, 70: 65-94.

Table 1: Regressions for Personal Income Across Japanese Prefectures.

Period	Basic Equation		Equations with District Dummies		Equations w/ Structural Variables + District Dummies	
	$\hat{\beta}$	$R^2[\hat{\sigma}]$	$\hat{\beta}$	$R^2[\hat{\sigma}]$	$\hat{\beta}$	$R^2[\hat{\sigma}]$
1930–1987	.0279 (.0031)	.92 [.0020]	.0280 (.0024)	.96 [.0013]		
1930–1955	.0358 (.0035)	.86 [.0045]	.0380 (.0037)	.90 [.0038]		
1955–1987	.0201 (.0036)	.57 [.0029]	.0230 (.0038)	.76 [.0022]		
1955–1960	−.0152 (.0079)	.07 [.0133]	−.0023 (.0082)	.34 [.0111]	.0047 (.0119)	.34 [.0112]
1960–1965	.0296 (.0072)	.30 [.0108]	.0360 (.0079)	.47 [.0093]	.0414 (.0096)	.47 [.0093]
1965–1970	−.0010 (.0062)	.00 [.0097]	.0128 (.0068)	.38 [.0076]	.0382 (.0091)	.54 [.0065]
1970–1975	.0967 (.0100)	.78 [.0095]	.0625 (.0091)	.85 [.0078]	.0661 (.0118)	.85 [.0079]
1975–1980	.0338 (.0100)	.23 [.0087]	.0455 (.0118)	.25 [.0085]	.0469 (.0145)	.23 [.0086]
1980–1987	−.0113 (.0059)	.07 [.0059]	.0029 (.0070)	.24 [.0053]	.0048 (.0073)	.23 [.0053]
Restricted	.0103 (.0037)	... ...	.0242 (.0037)	... ...	.0340 (.0044)	... ...
Likelihood Ratio	90.23		33.89		16.38	
p–Value	.0000		.0000		.0060	

Notes to Table 1: The regressions use non linear squares to estimate equations of the form:

$$(1/T)\ln(y_{it}/y_{i,t-T}) = a - [\ln(y_{i,t-T})](1-e^{\beta T})(1/T) + \text{other variables},$$

where $y_{i,t-T}$ is the per capita income in prefecture i at the beginning of the interval divided by the overall CPI. T is the length of the interval; 'other variables' are District dummies and structural variables (see description in the text).

Each column contains four numbers. The first one is the estimate of β. Underneath it (in parentheses) its standard error. To its right, the adjusted R^2 of the regression and below the R^2, the standard error of the equation. Thus, constant, District dummies and/or structural variables are not reported. The likelihood–ratio and p–values pertain to the test of the equality of the coefficients of the log of initial income over time. The p–value corresponds to a χ^2 with 5 degrees of freedom.

Table 2: Regressions for Personal Income Across U.S. States.

Period	Basic Equation		Equations with Regional Dummies		Equations w/ Structural Variables + Regional Dummies	
	$\hat{\beta}$	$R^2[\hat{\sigma}]$	$\hat{\beta}$	$R^2[\hat{\sigma}]$	$\hat{\beta}$	$R^2[\hat{\sigma}]$
1880–1988	.0171 (.0028)	.87 [.0016]	.0175 (.0046)	.92 [.0014]		
1880–1900	.0101 (.0022)	.36 [.0068]	.0224 (.0040)	.62 [.0054]	.0268 (.0048)	.65 [.0053]
1900–1920	.0218 (.0036)	.62 [.0029]	.0209 (.0038)	.67 [.0022]	.0269 (.0075)	.71 [.0060]
1920–1930	−.0149 (.0051)	.14 [.0132]	−.0122 (.0074)	.43 [.0111]	.0218 (.0112)	.64 [.0089]
1930–1940	.0141 (.0030)	.35 [.0073]	.0127 (.0051)	.36 [.0075]	.0119 (.0072)	.46 [.0071]
1940–1950	.0421 (.0048)	.72 [.0078]	.0373 (.0053)	.86 [.0057]	.0236 (.0060)	.89 [.0053]
1950–1960	.0190 (.0035)	.42 [.0050]	.0202 (.0052)	.49 [.0048]	.0305 (.0054)	.66 [.0041]
1960–1970	.0246 (.0039)	.51 [.0045]	.0135 (.0043)	.68 [.0037]	.1739 (.0053)	.72 [.0036]
1970–1980	.0198 (.0062)	.21 [.0060]	.0119 (.0069)	.36 [.0056]	.0042 (.0070)	.46 [.0052]
1980–1988	−.0060 (.0130)	.00 [.0142]	−.0005 (.0114)	.51 [.0103]	.0146 (.0099)	.76 [.0075]
Restricted	.0175 (.0013)	... ...	.0189 (.0019)	... ...	.0224 (.0022)	... ...
Likelihood Ratio	65.6		32.1		12.4	
p–Value	0.000		0.000		.134	

Notes to Table 2: The regressions use non linear squares to estimate equations of the form:

$$(1/T)\ln(y_{it}/y_{i,t-T}) = a - [\ln(y_{i,t-T})](1-e^{\beta T})(1/T) + \text{other variables,}$$

where $y_{i,t-T}$ is the per capita income in prefecture i at the beginning of the interval divided by the overall CPI. T is the length of the interval; 'other variables' are regional dummies and structural variables (see description in the text).

Each column contains four numbers. The first one is the estimate of β. Underneath it (in parenthesis) its standard error. To its right, the adjusted R^2 of the regression and below the R^2, the standard error of the equation. Thus, constant, regional dummies and/or structural variables are not reported. The likelihood–ratio and p–values pertain to the test of the equality of the coefficients of the log of initial income over time. The p–value corresponds to a χ^2 with 8 degrees of freedom.

Table 3: Regressions for Net Migration into Japanese Prefectures:
1955–1985

Period	Log of Personal Income	Own Density	Neighbors' Density	Extreme Temperature	R^2
1955–1987	.0173 (.0044)	−.0042 (.0017)	.0092 (.0016)	−.00001 (.00044)	.62
1955–1960	.0216 (.0036)	.0060 (.0013)	.0026 (.0019)	−.0001 (.0001)	.84
1960–1965	.0318 (.0058)	−.0019 (.0020)	.0148 (.0030)	−.0001 (.0001)	.72
1965–1970	.0345 (.0070)	−.0065 (.0017)	.0143 (.0025)	−.0001 (.0001)	.68
1970–1975	.0193 (.0060)	−.0063 (.0015)	.0114 (.0023)	−.0001 (.0001)	.48
1975–1980	.0061 (.0067)	−.0037 (.0011)	.0058 (.0014)	−.0001 (.0001)	.25
1980–1985	.0108 (.0043)	−.0021 (.0006)	.0044 (.0010)	−.0001 (.0001)	.38
Restricted	...	−.0027 (.0005)	.0062 (.0007)	−.0002 (.0002)	...
Restricted	.0271 (.0025)	−.0039 (.0005)	.0058 (.0008)	−.0004 (.0002)	...

Test of Stability of Y_0 coefficient: LR=65.27.

Notes to Table 3: The regressions use Iterative, Weighted Least Squares to estimate equations of the form:

$$m_{it} = a + b + \ln(y_{i,t-T}) + c_1 \, \text{Extreme}_i + c_2 \, \tau_{i,t-T} + c_3 \, \tau^{NE}_{i,t-T}$$

where m_{it} is the average annual net migration into prefecture i between years t–T and t, expressed as a ratio to the population at t–T. Extreme is a

measure of extreme temperature calculated as deviations of maximum and minimum temperatures from average temperature in prefecture i. $\tau_{i,t-T}$ is the population density (thousands of people per square kilometer) in state i at moment t–T. $\tau_{i,t-T}^{NE}$ is the population density of the neighboring prefectures (see text). Standard errors are in parentheses.

Table 4. Regressions for Net Migration into U.S. States: 1900–1987

Period	Log of Personal Income	Log of Heating Degree-Days	Density	Density^2	R^2
1900–1920	.0335 (.0075)	−.0066 (.0037)	−.0452 (.0077)	.0340 (.0092)	.70 [.0112]
1920–1930	.0363 (.0078)	−.0124 (.0027)	−.0452 (.0077)	.0340 (.0092)	.61 [.0079]
1930–1940	.0191 (.0037)	−.0048 (.0014)	−.0452 (.0077)	.0340 (.0092)	.71 [.0042]
1940–1950	.0262 (.0056)	−.0135 (.0022)	−.0452 (.0077)	.0340 (.0092)	.83 [.0065]
1950–1960	.0439 (.0085)	−.0205 (.0031)	−.0452 (.0077)	.0340 (.0092)	.76 [.0091]
1960–1970	.0436 (.0082)	−.0056 (.0025)	−.0452 (.0077)	.0340 (.0092)	.70 [.0069]
1970–1980	.0240 (.0091)	−.0076 (.0024)	−.0452 (.0077)	.0340 (.0092)	.73 [.0071]
1980–1987	.0177 (.0057)	−.0075 (.0018)	−.0452 (.0077)	.0340 (.0092)	.73 [.0049]
Restricted	.0261 (.0057)	...	−.0447 (.0078)	.0329 (.0093)	...

Notes to Table 4: The regressions use Iterative, Weighted Least Squares to estimate equations of the form:

$$m_{it} = a + b + \log(y_{i,t-T}) + c_1 \log(\text{Heat}_i) + c_2 \, \tau_{i,t-T} + c_3 \, \tau^2_{i,t-T}$$

where m_{it} is the average annual net migration into state i between years t–T and t expressed as a ratio of the population ar t–T. Heat_i is the average number of heating degree days for state i formed as an average of the cities in that state. $\tau_{i,t-T}$ is the population density (thousands of people per square mile) in state i at moment t–T. $\tau^2_{i,t-T}$ is the square of the population density. Standard errors are in parenthesis.

42

Figure 1
Convergence of Personal Income Across Japanese Prefectures
1930 Income and Income Growth 1930-1987

43
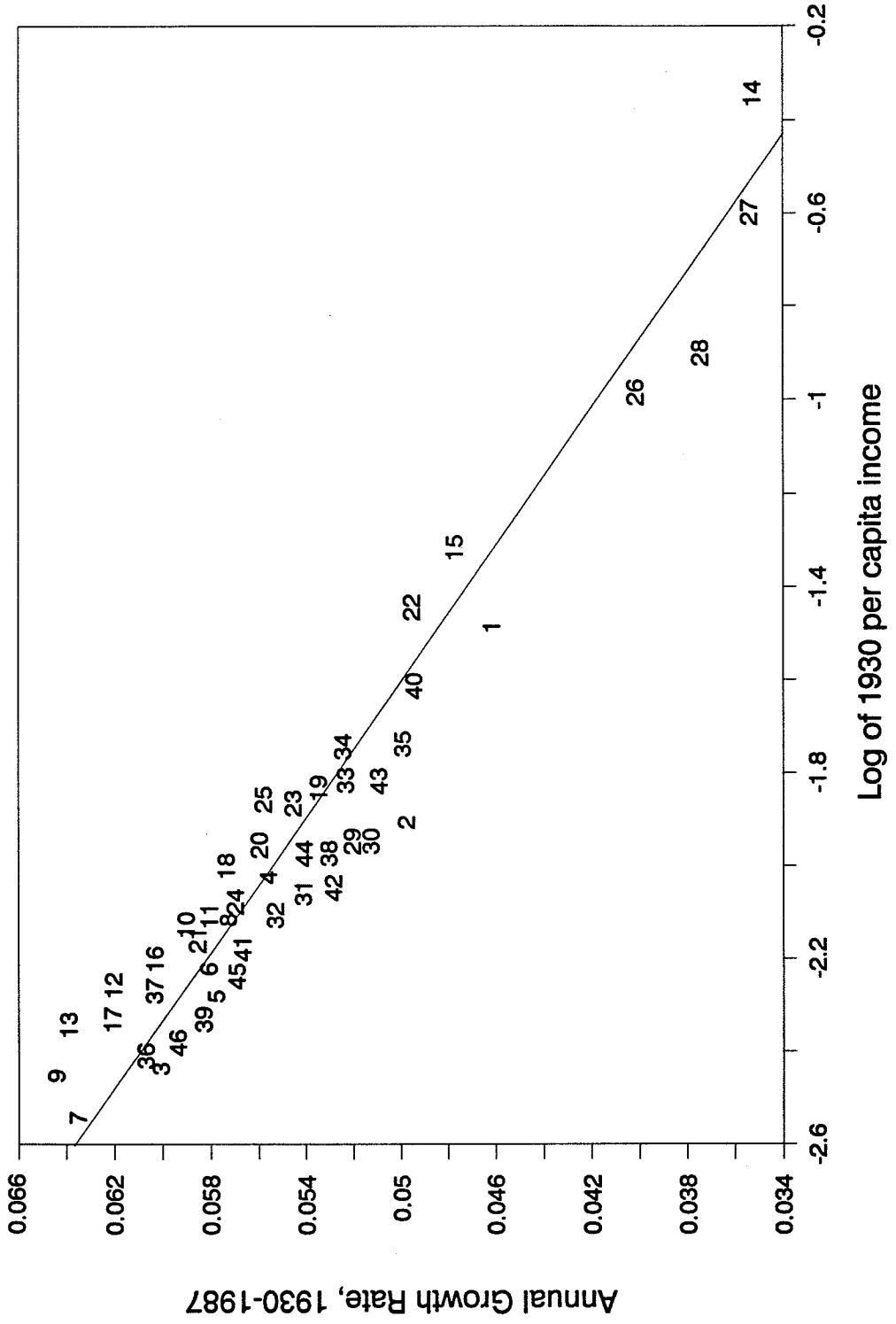
Figure 2

Convergence of Personal Income Across Japanese Districts
1930 Income and Income Growth 1930-1987
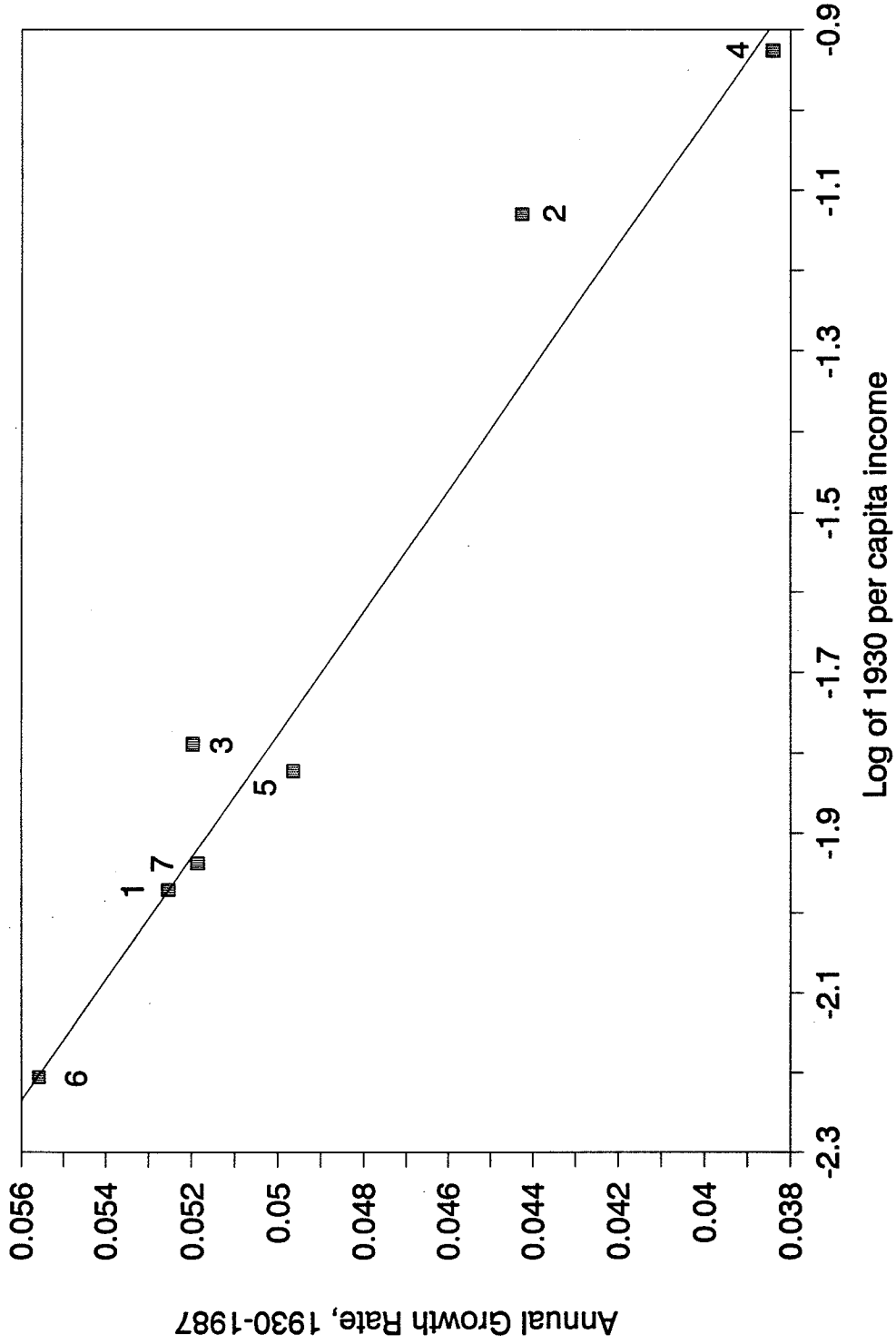
Figure 3
Convergence of Personal Income Across Japanese Prefectures
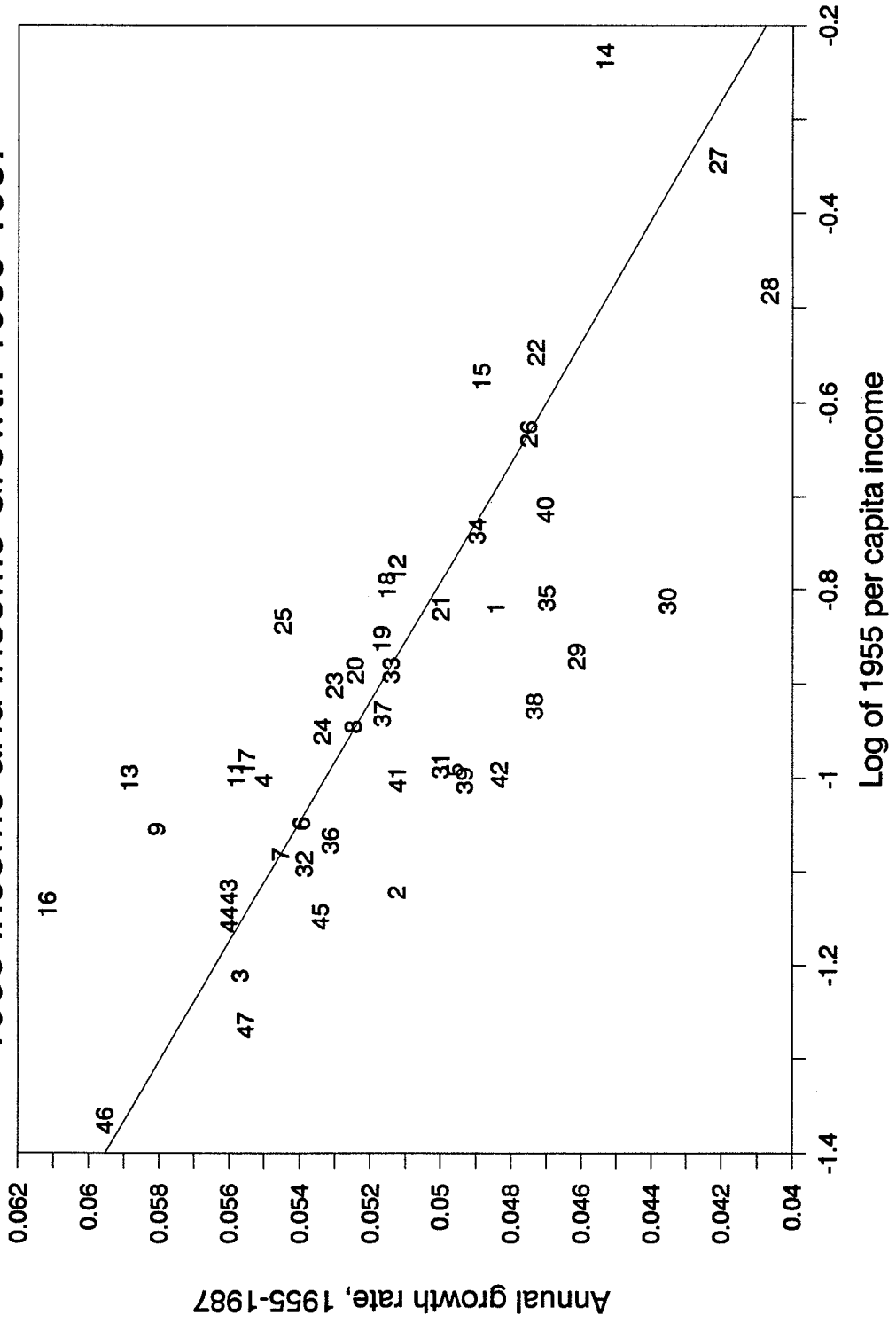1955 Income and Income Growth 1955-1987

Figure 4
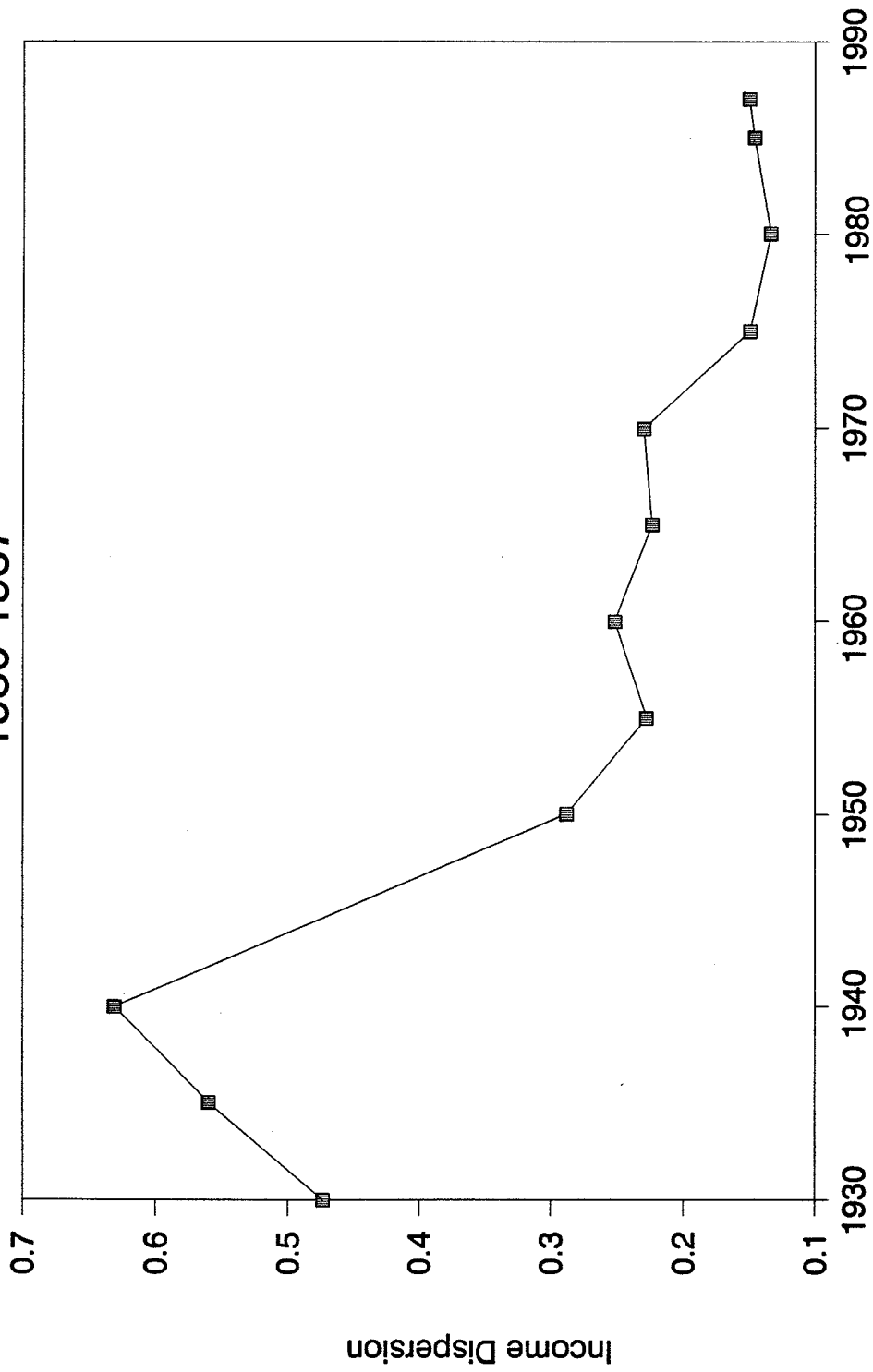Dispersion of Personal Income Across Japanese Prefectures
1930-1987

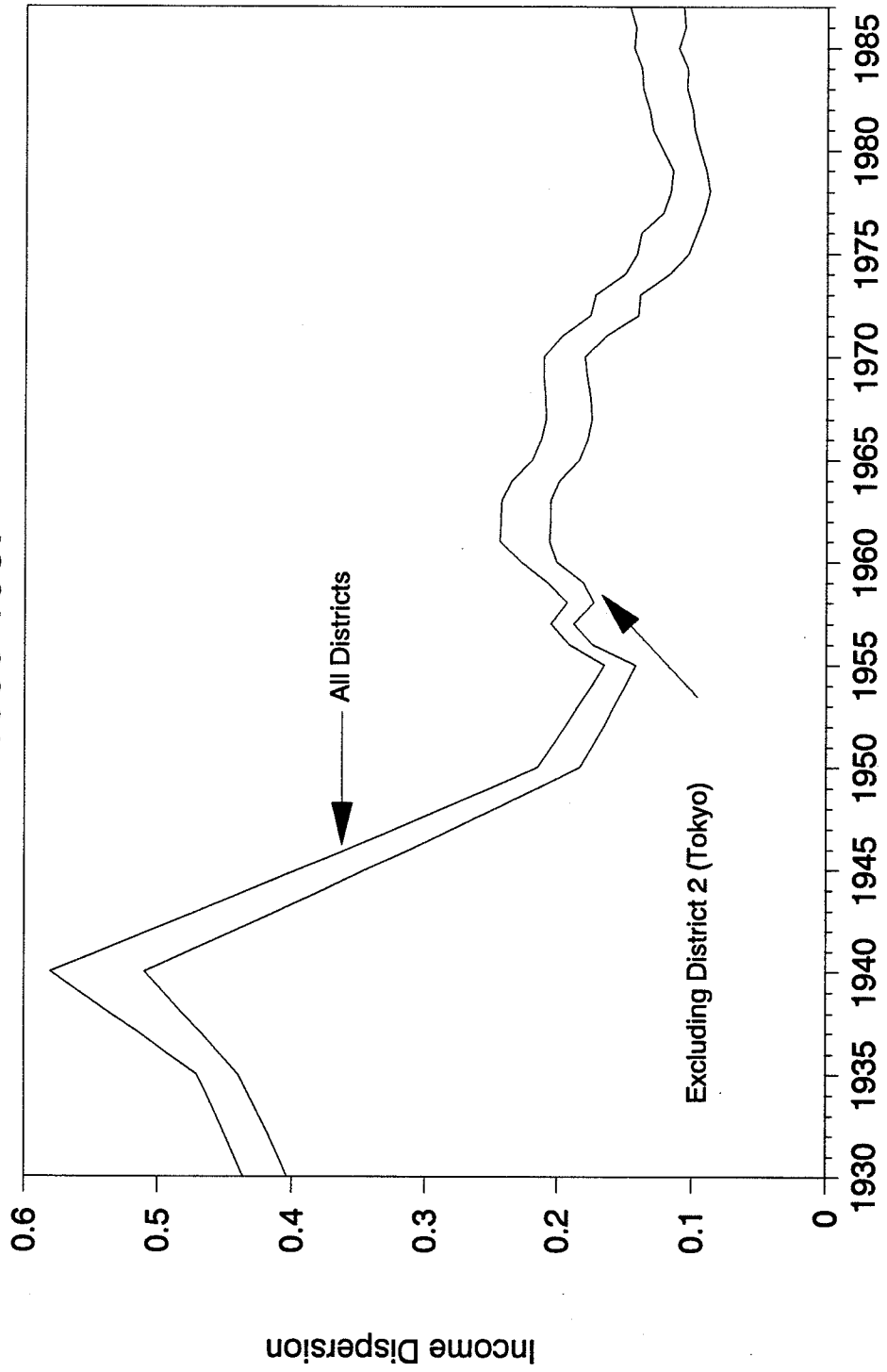Figure 5

Dispersion of Income Across Seven Japanese Districts
1930-1987

Income Dispersion

Figure 6

Convergence of Personal Income Across U.S. States
1880 Income and Income Growth 1880-1988
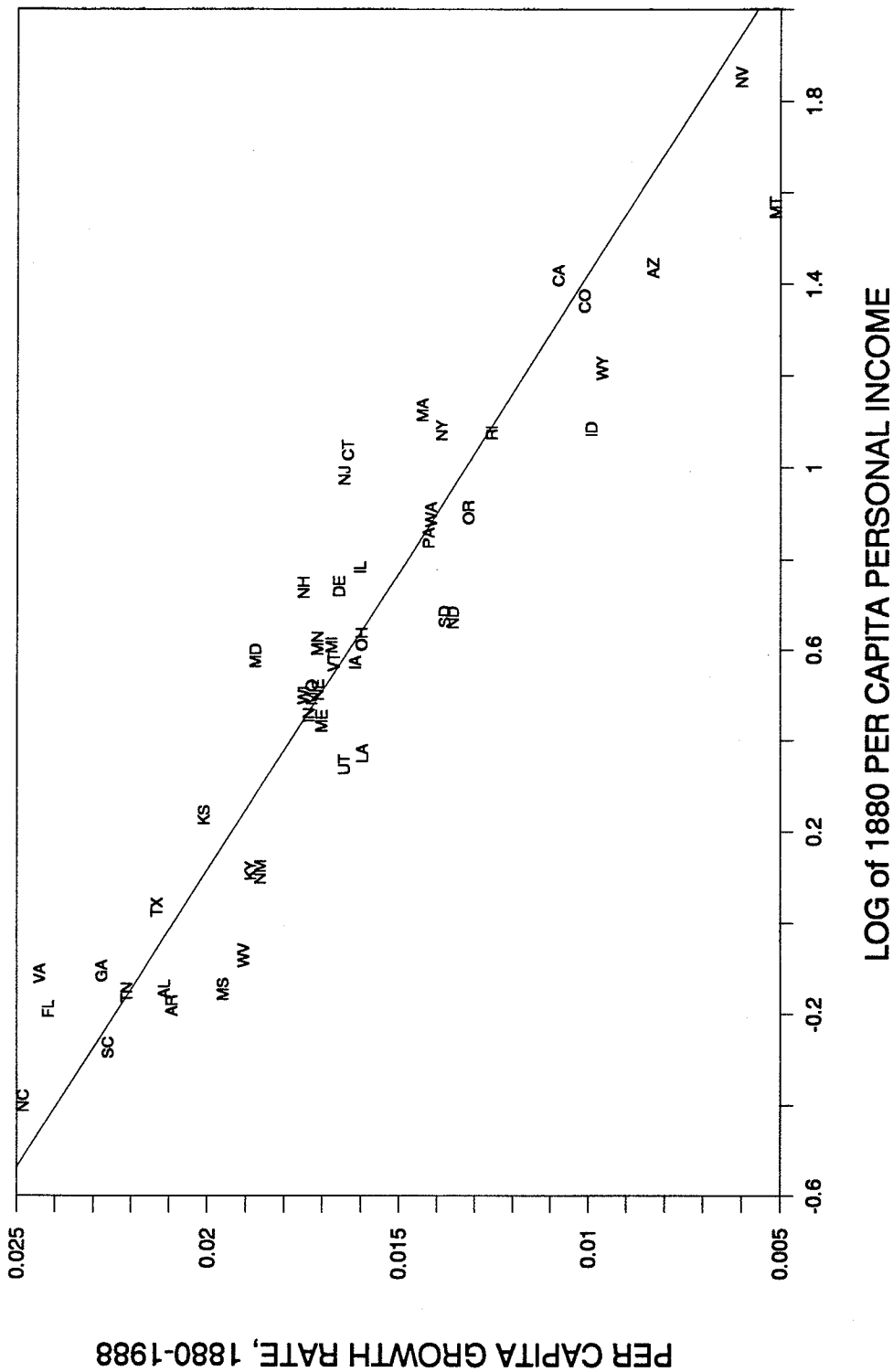
Figure 7
Convergence of Personal Income Across U.S. Regions
1880 Income and Growth 1880-1988
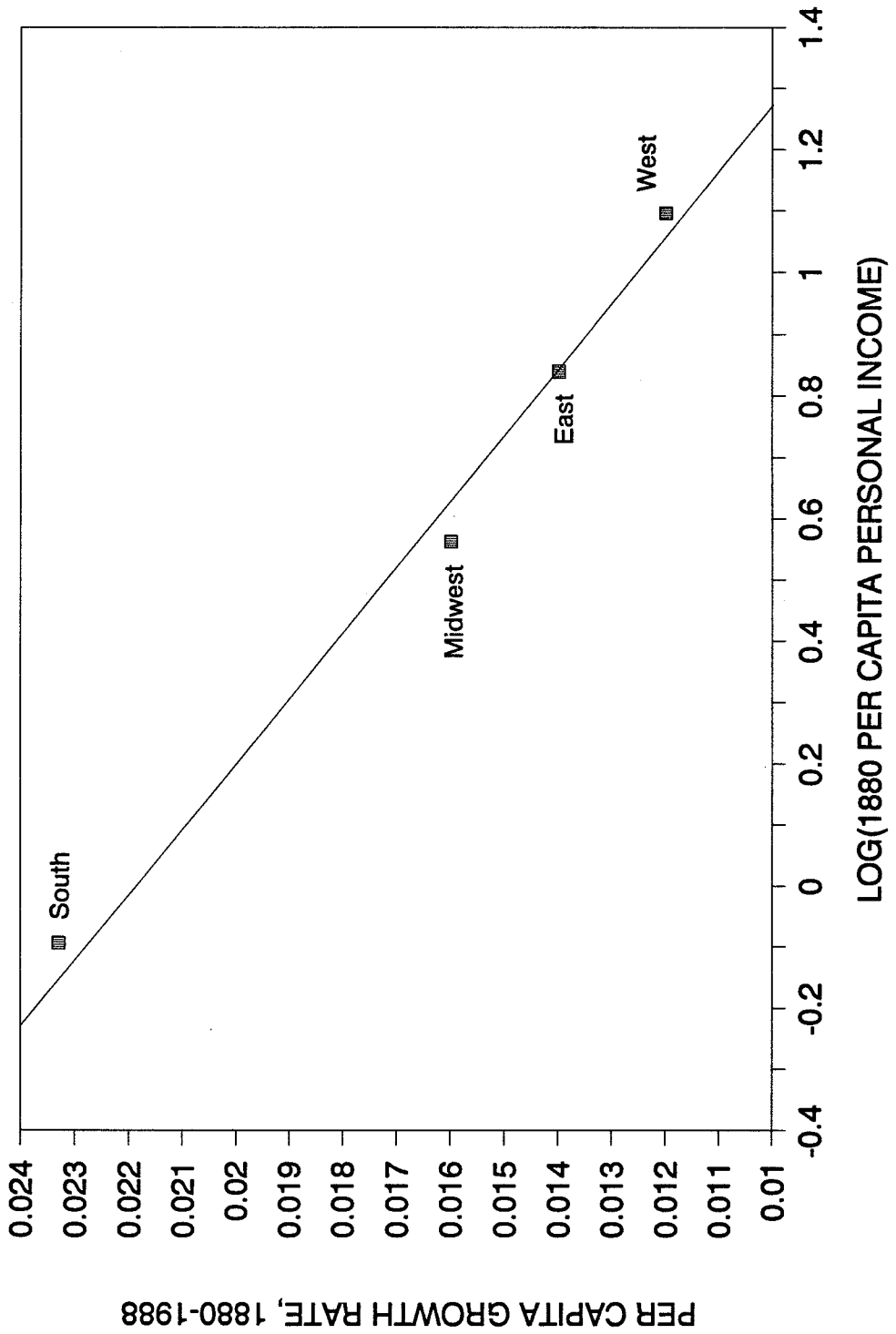
Figure 8

Dispersion of Personal Income Across U.S. States
1880-1988

50
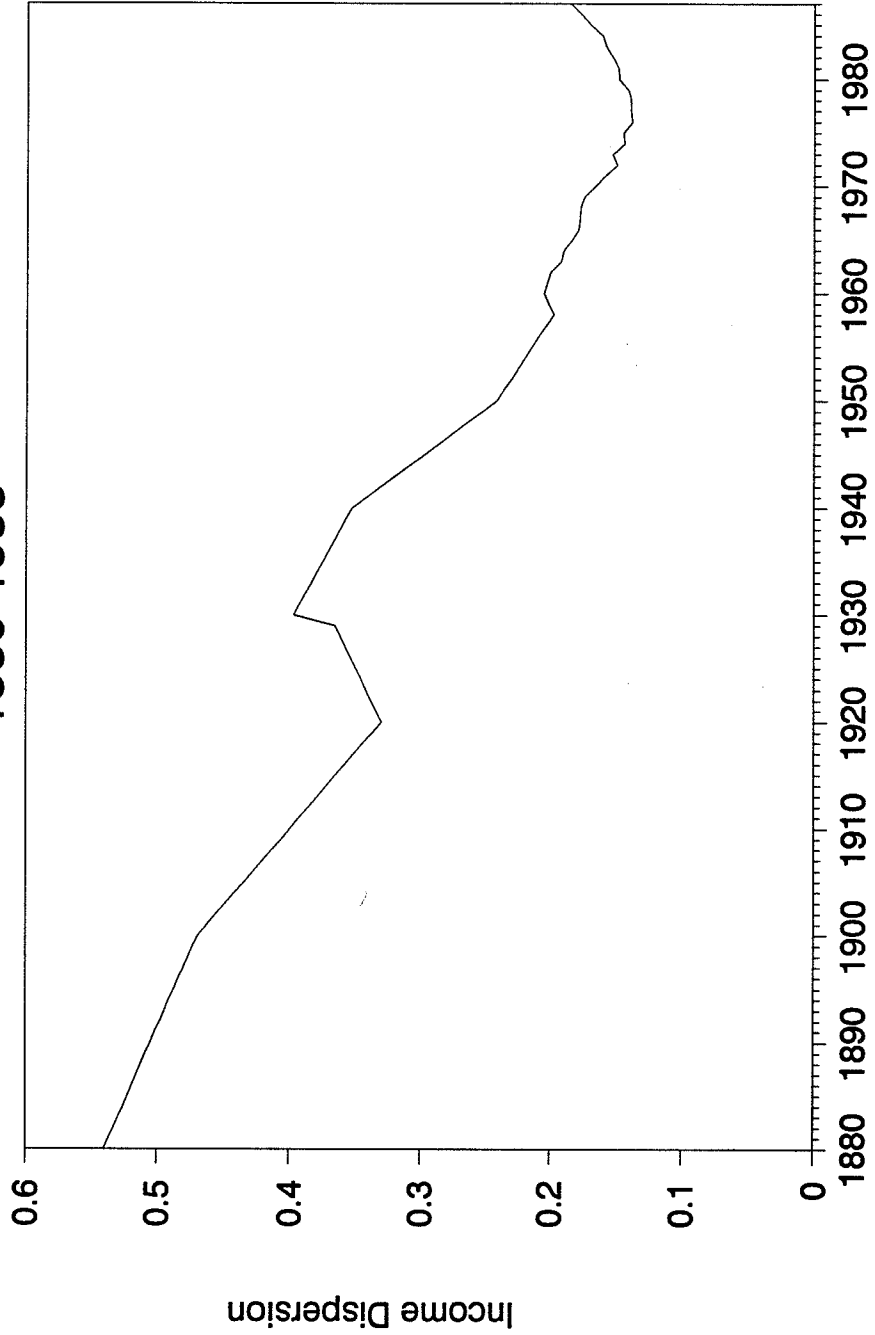
Figure 9
Dispersion of Personal Income Across Four U.S. Regions
1880-1988
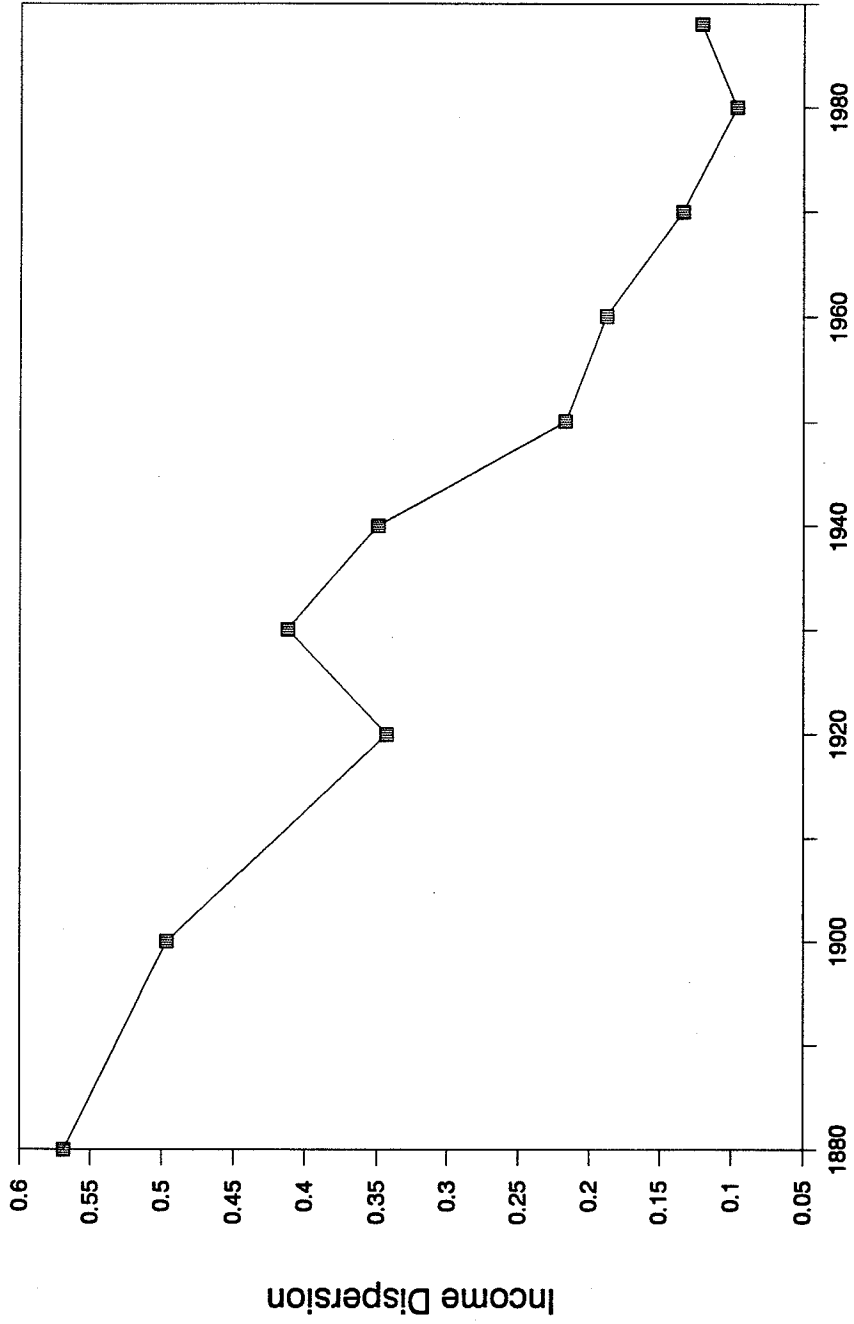
Figure 10

Migration and Initial Prefectural Income, 1955-1987
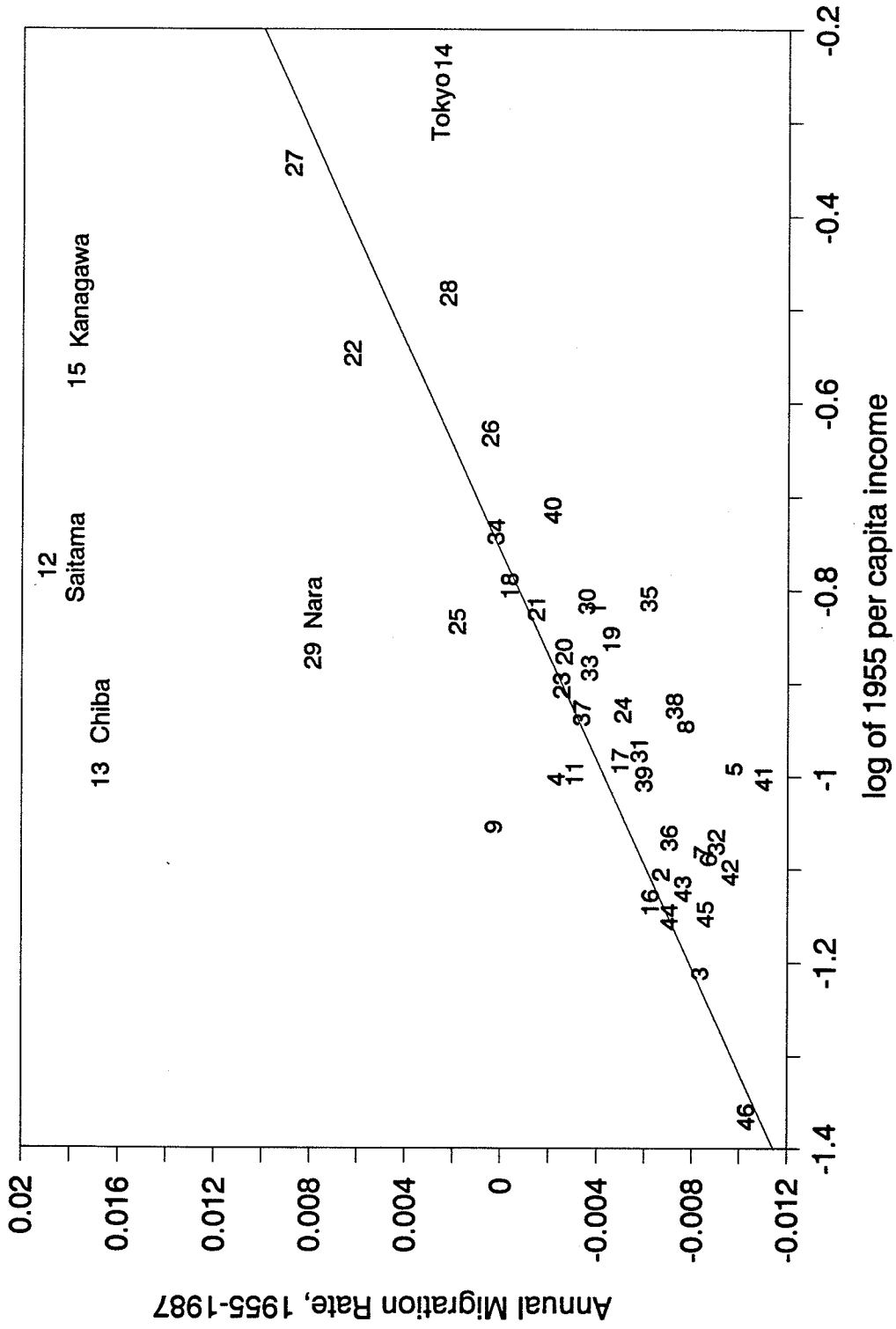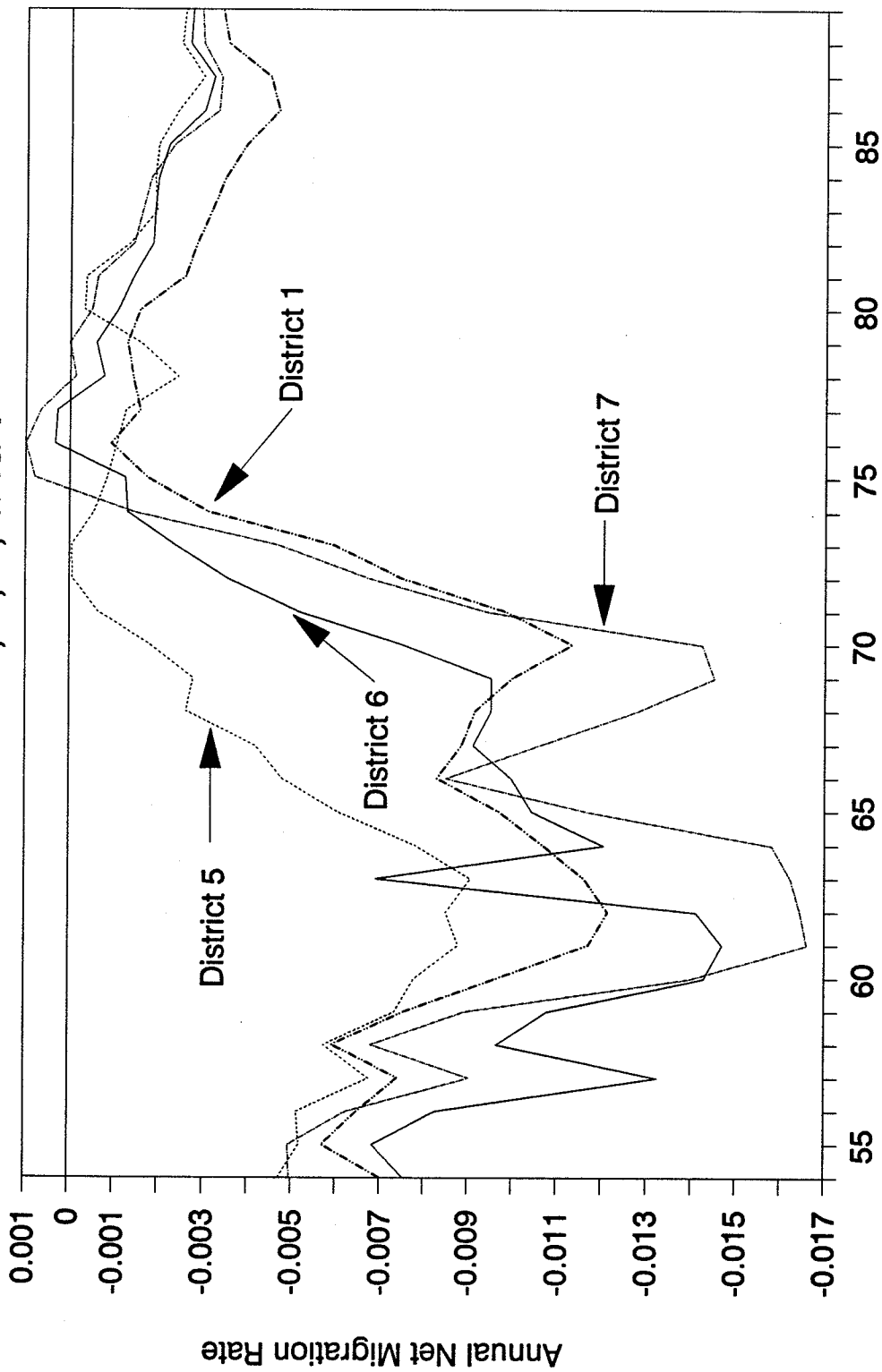
Figure 11a
Annual Net Migration Rates
Districts 1, 5, 6, and 7

Figure 11b
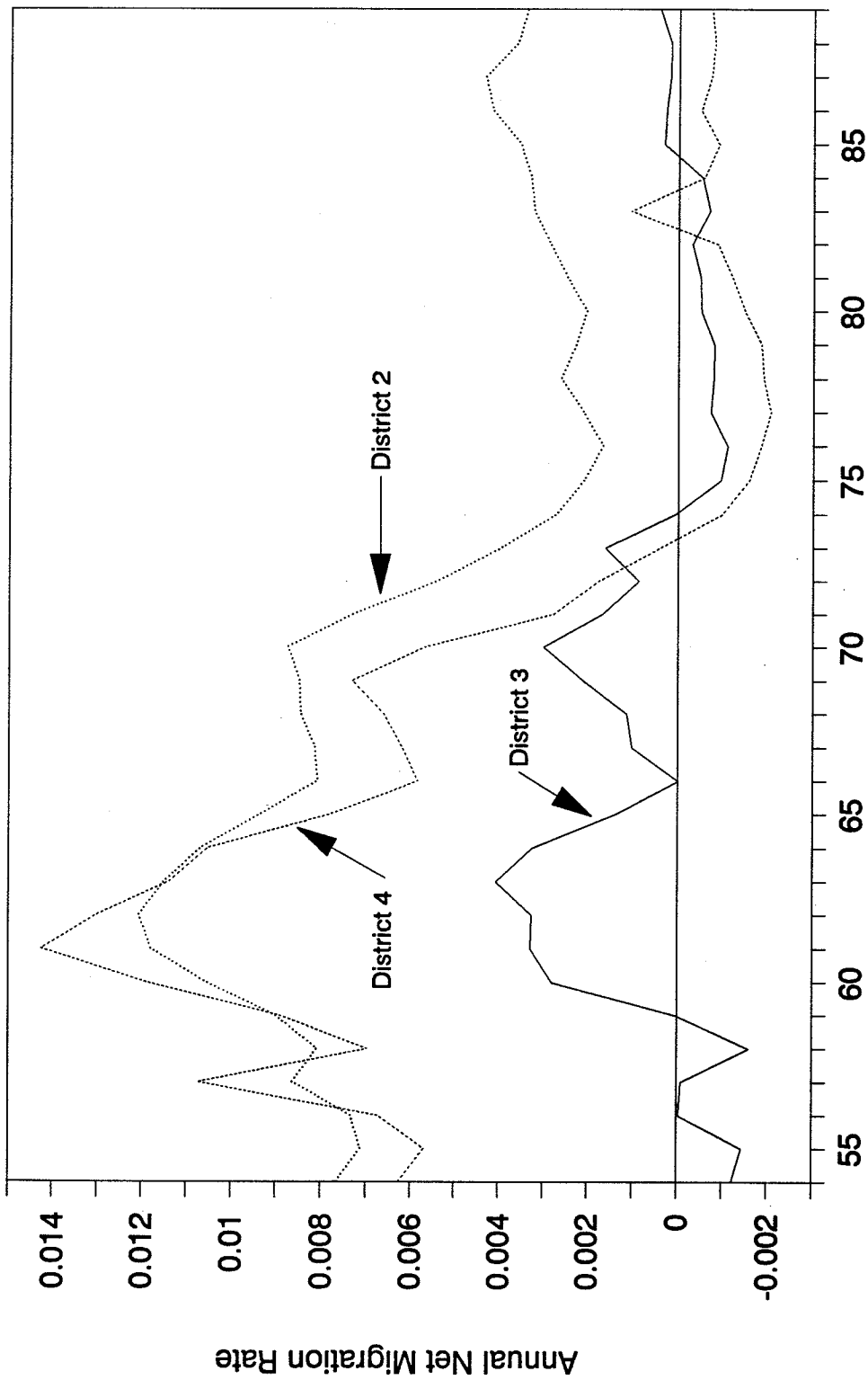Annual Net Migration Rates
Districts 2, 3 and 4

Figure 11c
Annual Net Migration Rates
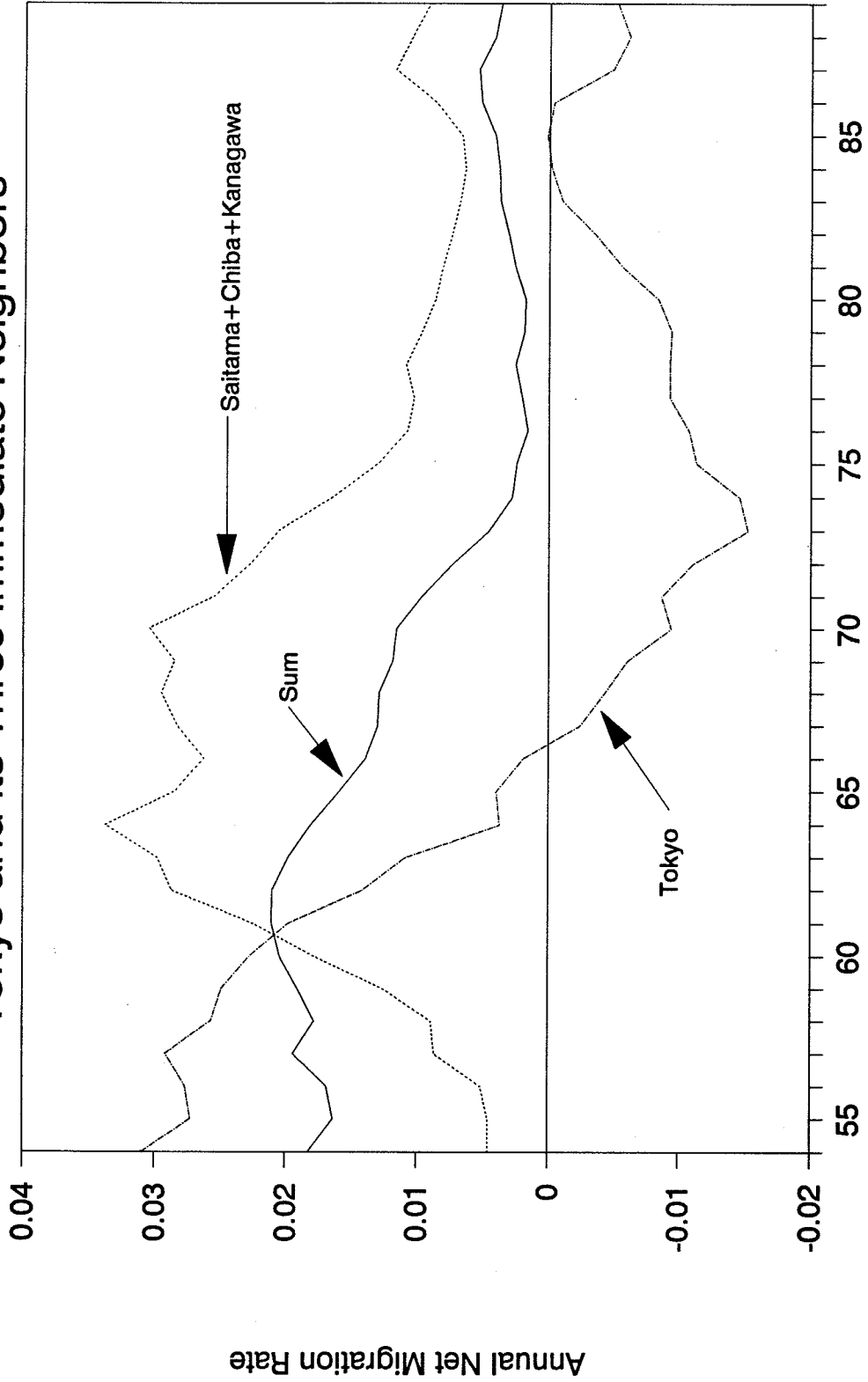Tokyo and its Three Immediate Neighbors

Figure 12

Migration and Initial State Income, 1900-1987