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Convergence across States and Regions

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Abstract

In this paper we examine the growth and dispersion of personal income in U.S. states and regions since 1880 and relate the patterns for individual states to the behavior of regions. Then we analyze the interplay between net migration and economic growth. We study the evolution of gross state product since 1963 and relate the behavior of aggregate product to productivity in eight major sectors. The overall evidence weighs heavily in favor of convergence: poor states tend to grow faster in terms of per capita income and product and within sectors as well as for state aggregates. The rate of convergence is, however, not rapid: the gap between the typical poor and rich state diminishes at roughly 2% per year.

We apply the same framework to patterns of convergence across 73 regions of seven European countries since 1950. The process of convergence within European countries is similar to that for the United States. In particular, the rate of convergence is again about 2% per year.

KEY WORDS: Convergence, Growth, Migration, Regional Economics

An important economic question is whether poor countries or regions tend to converge toward rich ones. We want to know, for example, whether the poor countries of Africa, South Asia, and Latin America will grow faster than the developed countries, whether southern Italy will become like the north, whether and how fast the eastern regions of Germany will attain the prosperity of the western regions, and—in an historical context—how the American south became nearly as well off as the north.

Although some economic theories predict convergence, the empirical evidence has been a subject of debate. We add to the evidence in this study by extending our previous analysis of economic growth across the U.S. states (Barro and Sala–i–Martin [1990]). We examine the growth and dispersion of personal income since 1880 and relate the patterns for individual states to the behavior of regions. Then we analyze the interplay between net migration and economic growth. We study the evolution of gross state product since 1963 and relate the behavior of aggregate product to productivity in eight major sectors. The overall evidence weighs heavily in favor of convergence: poor states tend to grow faster in terms of per capita income and product and within sectors as well as for state aggregates. The rate of convergence is, however, not rapid: the gap between the typical poor and rich state diminishes at roughly 2% per year.

We apply the same framework to patterns of convergence across 73 regions of 7 European countries since 1950. The process of convergence within the European countries is in many respects similar to that for the United States. In particular, the rate of convergence is again about 2% per year.

We conclude by using the findings to forecast the convergence process for the eastern regions of unified Germany. The results are not very encouraging: if the histories of the U.S. states and European regions are useful guides, then the convergence process will occur, but only at a slow pace.

Framework of the Analysis

Our previous study (Barro and Sala–i–Martin [1990], referred to henceforth as B/X) studied convergence patterns for economic growth across the U.S. states. We based the main analysis on a growth equation that derives, as a log–linear approximation, from the transition path of the neoclassical growth model for closed economies (Solow [1956], Cass [1965], Koopmans [1965]). We follow this research strategy again in this paper; that is, we begin with the closed–economy framework and then consider how the model would be affected by open–economy elements, which would be important for the U.S. states and European regions.

We showed in B/X that the transitional growth process in the neoclassical model can be approximated as

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

where i indexes the economy and t indexes time, y_{it} is per capita output (equal to income per person and per worker in the standard model), x_i is the steady–state per capita growth rate (corresponding to exogenous, labor–augmenting technological progress in the standard model), $\hat{y}_{it}$ is output per effective worker (that is, the number of workers adjusted for the effect of technological progress), $\hat{y}_i^*$ is the steady–state level of output per effective worker, T is the length of the observation interval, β is the speed of convergence, and u_{it} is an error term. (The error term is a distributed lag of disturbances between dates t–T and t.) The coefficient β indicates the rate at which $\hat{y}_{it}$ approaches $\hat{y}_i^*$.

On the production side, the neoclassical model assumes diminishing returns to capital, exogenous technological progress, full employment, a fixed relation between the labor force and population, and exogenous growth of population. With respect to preferences, the model assumes that the saving rate derives from choices of utility–maximizing households over an infinite horizon. (The infinite horizon can represent finite–lived individuals who are connected to their descendants via a chain of intergenerational transfers.) The steady–state value, $\hat{y}_i^*$, depends on the parameters of technology and preferences. We can extend the notion of technology to include natural resources, such as geography, fertile land, and the availability of minerals, and to (exogenous) government policies that affect property rights, the provision of infrastructure services, tax rates, and so on.

The convergence coefficient β depends on the productivity of capital and the willingness to save. In particular, the source of convergence in the neoclassical growth model is the assumed diminishing returns to capital. If the ratio of capital (and hence, output) to effective labor is further below the steady–state ratio, then the marginal product of capital is higher. Therefore, for given saving behavior, an economy grows faster the further it is below the steady state, that is, the higher $\hat{y}_i^*/\hat{y}_{i,t-T}$ in equation (1). If we compare different production functions, then β is higher if diminishing returns to capital set in faster. For example, for a Cobb–Douglas production function with capital share α, a smaller α corresponds to a larger β. As α tends to one, so that diminishing returns to capital no longer apply, the convergence coefficient β tends to zero. This case corresponds to endogenous growth models with constant returns to a broad concept of capital, as discussed in Rebelo (1990).

The parameter β also depends on saving behavior, although more on the variation in the saving rate over the transition to the steady state than on the level of the saving rate. We explore these effects in Barro and Sala–i–Martin (1991, Ch. 1). One result is that a greater willingness to substitute intertemporally tends to raise β.

Although the coefficient β can differ across economies, we neglect these differences in the subsequent discussion. This assumption is probably satisfactory for the U.S.

states, which are likely to be similar in terms of the underlying parameters of technology and preferences. Also, the theory implies that pure differences in the level of technology—and hence, the spread in levels of per capita income that derives from these technological differences—do not affect β. Thus, the convergence coefficients β can be similar across economies that appear in other respects to be very different.

We noted in B/X that our empirical estimates of β for the U.S. states—somewhat greater than 2% per year—accord with the neoclassical growth model only if diminishing returns to capital set in very slowly. For example, with a Cobb–Douglas production function, the capital–share coefficient, α, has to be in the neighborhood of 0.8. We also argued that this high value for α is reasonable if we take the appropriately broad view of capital to include non–human and human components. That is, education and other expenditures on people are important parts of the investment process.

The closed–economy model cannot, of course, be applied literally to the U.S. states or the regions of other countries. We discussed in B/X some implications of capital mobility. If technologies are the same, then convergence in per capita outputs and capital stocks tends to occur more rapidly than in the closed–economy setting, whereas convergence in per capita incomes and assets tends to occur less rapidly. Models that assume perfect capital mobility tend to have counterfactual implications, such as the prediction that the most patient economy owns everything asymptotically and that less patient economies eventually have negative assets and negligible consumption per effective worker.[1] Also, sharp distinctions in the behavior of output and income do not show up in our empirical findings. We have considered models in which "imperfections" in capital markets imply that only a fraction of physical capital serves as collateral on loans (see Cohen and Sachs (1986) and Barro and Sala–i–Martin [1991, Ch. 2]). In the context of the U.S. states, this setting applies if the residents or government of a state cannot borrow nationally to finance all of their expenditures on education and other forms of investment in human capital. This type of model predicts that product and income eventually behave in a similar manner and that each exhibits the kind of convergence property implied by the closed–economy specification of equation (1).

We have extended the framework to allow for the mobility of labor. Raw labor tends to migrate toward richer economies, which have higher wage rates. This movement of persons lowers the capital–labor ratio in places with initially high ratios; hence, diminishing returns to capital set in more rapidly and the convergence coefficient β in equation (1) is higher for given parameters of preferences and technology.[2] In other words, for given values of the other parameters, the capital–share coefficient α for a Cobb–Douglas production function would have to be even higher than 0.8 to be consistent with the empirical estimates of β.

The rate of convergence tends also to be higher if we allow for the flow of technological advances from rich to poor economies (see Nelson and Phelps [1966] for an early model of technological diffusion.) Differences in levels of technology tend, however, to alter the implications of capital mobility. Human and physical capital may move from poor to rich economies and thereby create a force toward divergence.

We discussed in B/X two concepts of convergence related to equation (1). The first, called β convergence, relates to poor economies growing faster than rich ones, and the second, called σ convergence, involves a decline over time in the cross–sectional dispersion of per capita income or product. The model implies a form of ***conditional*** ***β*** convergence in that, for given steady–state values x_i and $\hat{y}_i^*$, an economy's per capita growth rate is higher the lower the starting level of per capita output, $y_{i,t-T}$. The convergence is conditional in that $\hat{y}_{i,t-T}$ enters in relation to $\hat{y}_i^*$, which may differ across economies. The coefficient β measures the speed of this conditional convergence. To isolate β empirically we have to hold constant differences in the steady–state values x_i and $\hat{y}_i^*$. Although we included some additional variables as proxies for these differences, we found that the variations in the steady–state values seemed to be minor for the U.S. states. These variations appeared to be more significant for a group of relatively homogeneous countries, such as the OECD members, and were considerably more important for a broad sample of 98 countries.

Even if x_i and $\hat{y}_i^*$ are identical for a group of economies, a positive β coefficient need not imply that the cross–sectional dispersion of per capita output, y_{it}, diminishes over time. The positive β tends to reduce the dispersion in $\log(y_{it})$ from equation (1), but new shocks u_{it} tend to raise it. Equation (1) implies, for a given distribution of u_{it}, that the cross–sectional standard deviation of $\log(y_{it})$, denoted σ_t, approaches a constant σ. The dispersion σ_t falls (or rises) over time if it starts above (or below) σ; hence, β convergence (in the sense of $\beta>0$) need not imply σ convergence (in the sense of a declining σ_t). If the steady–state values $\hat{y}_i^*$ differ across economies, then we could also consider a conditional form of σ convergence. That is, conditional σ convergence applies if the dispersion of the deviations, $\log(\hat{y}_{it})-\log(\hat{y}_i^*)$, diminishes over time. Because this concept relies heavily on measures of the $\hat{y}_i^*$ we have not attempted to implement this idea.

We can bring out the distinction between β and σ convergence by considering two different kinds of questions. Suppose that we are interested in how fast and to what extent the per capita income of a particular economy (or person) is likely to catch up to the average of per capita incomes across economies (or persons). Then β convergence is the concept that matters for the answer. Suppose, on the other hand, that we want to know how the distribution of per capita income across economies (or persons) has behaved in the past or is likely to behave in the future. In this case, σ convergence is the relevant concept.[3]

Many disturbances, such as war or shocks to agriculture or oil, affect economies differentially. These disturbances tend to raise the cross–sectional variance of u_{it}

temporarily and thereby raise σ_t above σ. Subsequently, if the long–run distribution of u_{it} is unchanged, σ_t tends to fall gradually back to the value σ. Events related to war, agriculture, oil, etc., also affect groups of economies in a correlated manner; contingent on one of these events, the u_{it} are not independent draws over the economies i. We used regional dummies and measures of sectoral composition of output to address this problem in our previous work. That is, we treated the u_{it} as independent over i once we included these additional variables in the regressions.

The failure to introduce these additional variables can also lead to biased estimates of the coefficient β, contingent on the realization of a particular shock. Consider, for example, an adverse shock to agricultural output in a setting in which agricultural economies start with below average per capita product. Because of the positive correlation of the shock with $y_{i,t-T}$, we would underestimate β if we did not hold constant the shock.

Personal Income across the U.S. States

Figure 1 shows the broad pattern of β convergence for per capita personal income, exclusive of all transfers, for 47 U.S. states or territories from 1880 to 1988.[4] The figure shows the strong negative correlation (–0.93) between the average growth rate from 1880 to 1988 and the log of per capita personal income in 1880. The means and standard deviations for these and other variables are in Appendix Table 1.

One aspect of Figure 1 is that the southern states tend to have low per capita income in 1880 and high average growth rates thereafter. Less well known is that the western states have above–average per capita income in 1880 and below–average growth rates afterwards.

Although regional catch–up is part of the overall convergence story, Figures 2 and 3 show that the pattern between regions (east, south, midwest, west) is similar to that within regions. Figure 2 shows the four data points that correspond to the regional means. Figure 3 shows the pattern when the state growth rates and logs of initial income are measured relative to the respective regional means. The relations between growth rates and starting levels from Figures 2 and 3 are quantitatively similar.

Table 1 shows regression estimates of the convergence coefficient β for nine sub–periods of the sample from 1880 to 1988. The first column is a form of equation (1) that includes for each sub–period only a constant and the log of the state's initial personal income per capita. Although most of the estimated coefficients are significantly positive, the magnitudes vary a great deal and two of the point estimates are negative. (Recall that we define the coefficient β in equation (1) so that a positive β means that poor economies grow faster than rich ones.) If we constrain the estimate of β to be the same for all sub–periods, then the resulting joint estimate is $\hat{\beta}$=0.0175 (s.e.=0.0013), that is, in the neighborhood of 2% per year. However, a likelihood–ratio test, shown in the table, strongly rejects the hypothesis that β is stable over the sub–periods (p–value = 0.000).[5]

Column 2 of Table 1 introduces regional dummies (hence, three new variables) for each sub–period. These dummies proxy for differences in the steady–state values x_i and $\hat{y}_i^*$ in equation (1) and also absorb fixed effects related to regions in the error term u_{it}. Although the fits are improved and the variation in $\hat{\beta}$ over sub–periods is somewhat reduced, the results still reject the hypothesis of stability in β over the periods (p–value = 0.000). The restricted point estimate, $\hat{\beta}$, for the nine sub–periods, 0.0189 (0.0019), is similar to that reported in column 1. The estimate from column 2 reflects within–region β convergence, whereas that from column 1 reflects a combination of within– and between–region convergence. Hence, as also suggested by Figures 1–3, the results indicate that the within– and between–region rates of β convergence are similar.

Column 3 of Table 1 adds two additional variables. The first, denoted $AGRY_{i,t-T}$, is the share of income originating in agriculture in state i's personal income at the start of each sub–period (that is, in year t–T). This variable is available for all of the sub–periods since 1880. As with the regional dummies, the agriculture variable can hold constant differences in steady–state values, x_i and $\hat{y}_i^*$ in equation (1), as well as common effects related to agriculture in the error term u_{it}.

The second variable, denoted S_{it} for structure, relates to the breakdown of state i's personal income into nine standard sectors: agriculture, mining, construction, manufacturing, transportation, wholesale and retail trade, finance–insurance–real estate, services, and government. We first compute the national growth rates of income originating in each sector for each sub–period (less the growth rate of national population for the period). Then we weight the national growth rates by the share of each sector in state i's personal income at the start of the sub–period (that is, in year t–T). Hence, the formula for S_{it} is

$$(2) \qquad S_{it} = \sum_{j=1}^{9} w_{ij,t-T} \cdot \log(y_{jt}/y_{j,t-T})$$

where $w_{ij,t-T}$ 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 in sector j at time t, expressed as a ratio to national population at time t. Aside from the effect of changing sectoral weights within a state, the variable S_{it} would equal the growth rate of per capita personal income in state i between years t–T and t if each of the state's sectors grew at the national average rate for that sector. In particular, the variable S_{it} reflects shocks to agriculture, oil, etc., in a way that interacts with state i's concentration in the sectors that do relatively well or badly because of the shocks. We think of S_{it} as a proxy for common effects related to sectoral composition in the error term u_{it} in equation (1). Note that S_{it} depends on contemporaneous realizations of national variables, but only on lagged values of state variables. Because the impact of an individual state on national aggregates is small, S_{it} can be (nearly) exogenous with respect to the individual error term for state i.

We have the data to construct S_{it} only since 1929. For that reason, we also include the variable $AGRY_{i,t-T}$, described above, as a separate influence. We include $AGRY_{i,t-T}$ for all sub–periods, although the results are similar if we omit this variable for the sub–periods beginning after 1929.

Column 3 of Table 1 shows the estimates for β when the two variables, $AGRY_{i,t-T}$ and S_{it}, are included in the regressions. The principal new finding is that the estimates $\hat{\beta}$ are much more stable across periods. The greater stability arises because we hold constant shocks for some sub–periods that are correlated with initial per capita income and also affect groups of states in common. For example, agriculture suffered relative to other sectors in the 1920s.[6] Because agricultural states had below–average per capita income in 1920, we estimate negative β coefficients for the 1920–1930 sub–period in columns 1 and 2 of Table 1. But, once we hold constant the differences in agricultural shares, $AGRY_{i,t-T}$, in column 3, we estimate a β coefficient for this sub–period, 0.0218 (0.0112), that is similar to those found for the other sub–periods. The joint estimate for the nine sub–periods is now $\hat{\beta}$=0.0224 (0.0022) and we accept the hypothesis of coefficient stability at the 5% level (p–value = 0.13). Thus, as noted before, the results suggest β convergence at a rate somewhat above 2% per year.[7]

The estimates of β convergence shown in column 3 of the table are net of compositional effects from shifts of persons out of agriculture and toward higher productivity jobs in industry and services.[8] These effects are held constant by the initial agricultural shares, $AGRY_{i,t-T}$, which are included as regressors. In particular, if we add the change in the agricultural share, $AGRY_{it}-AGRY_{i,t-T}$, to the regressions, then the joint estimate of β is virtually unchanged from that shown in column 3. In general, industry–mix effects matter for the results if shifts of income shares among sectors with different average levels of productivity are correlated with initial levels of per capita income. It is unclear that we would want to filter out this kind of effect in measuring β convergence, but, in any event, our examination of productivity data from the post–World War II period suggests that shifts between agriculture and non-agriculture would be the principal effect of this type.

We also have regression estimates that parallel those in column 3 but exploit only the between–region variation in growth rates. Because we have 4 regions and 9 sub–periods, we now have 36 observations of per capita growth rates. (With a single β coefficient, this system has 25 independent variables and therefore 11 degrees of freedom.) The joint estimate of β is 0.0187 (0.0069), which does not differ greatly from the joint estimate shown in column 3. Thus, as noted before, the between-region β convergence is similar to the within–region convergence. We would not get this similarity if the states in the four regions differed substantially (after holding constant regional differences in the variables $AGRY_{i,t-T}$ and S_{it}) in terms of the steady–state values x_i and $\hat{y}_i^*$ in equation (1). Thus, the findings suggest that the regions are converging toward similar steady–state behavior of per capita income.

Figure 4 shows the (unweighted) cross–sectional standard deviation, σ_t, for the log of per capita personal income for 48 U.S. states from 1880 to 1988. (The observation for 1880 applies to 47 states or territories. The data are plotted for 1880, 1900, 1920, 1929, 1930, 1940, 1950, and annually since 1955.) We concentrate for now on the data without transfers (from all levels of government), which are the figures that we have used thus far.

Figure 4 shows that σ_t declined from 0.54 in 1880 to 0.33 in 1920, but then rose to 0.40 in 1930. This increase reflects the adverse shock to agriculture during the 1920s; the effect on σ_t is pronounced because the disturbance adversely affected states that were already below average in per capita income. After 1930, σ_t fell to 0.35 in 1940, 0.24 in 1950, 0.21 in 1960, 0.17 in 1970, and a low point of 0.14 in 1976. The sharp decline during the 1940s reflects the favorable experience of agriculture.[9] The pattern of long–term decline in σ_t reversed after the mid–1970s and σ_t rose to 0.15 in 1980 and 0.19 in 1988. We think that the increase in σ_t after the mid–1970s relates to oil shocks. A later section discusses these effects in the context of a comparison of results for personal income with those based on gross state product.

The broad observation from Figure 4 is a long–term decline in σ_t from a value above 0.5 to a plateau around 0.15–0.20. This pattern accords with the σ convergence predicted by the neoclassical growth model if the states began in 1880 with a dispersion that was well above the steady–state amount, σ. If we use the observed values, $\sigma_{1880} =$ 0.545 and $\sigma_{1988} = 0.194$, and the previous estimate, $\beta = 0.02$ per year, then we can estimate the standard deviation, σ_u, of the (annual) error term u_{it} in equation (1) as 0.037. This value implies that the steady–state dispersion, σ, is 0.18.[10]

One aspect of the high dispersion in 1880 is the low per capita incomes of southern states relative to those of non–southern states, a pattern that can be traced back to the Civil War. As discussed in B/X, the average income in the south was not very much below that in the non–south in 1840, but was about 50% of that in the non–south in 1880. Another element is that the western states, which were in many respects new territories in 1880, had relatively high per capita incomes at the start of the sample. Some of this high income represented temporary opportunities in mining.

Figure 4 also shows the values of σ_t computed from personal income inclusive of transfer payments.[11] (Our data on transfers do not separate the amounts received from the federal government from those received from state and local governments.) The inclusion of transfers lowers the level of σ_t because the ratio of transfers to personal income exclusive of transfers is substantially negatively correlated with the log of personal income exclusive of transfers. For example, the correlation between the transfer ratio and the log of personal income exclusive of transfers is –0.76 in 1987. On the other hand, the time pattern for σ_t inclusive of transfers is similar to that exclusive of transfers. The quantitative effect of the transfer component has been increasing over time. Although the effect in 1950 is negligible, in 1987, σ_t exclusive of transfers is 0.187, whereas that inclusive of transfers is 0.165.

We now return to the personal income data measured exclusive of transfers. Figure 5 shows the value σ_t when computed only from the four regional averages for each date. The main observation is that the regional pattern in σ_t pretty well matches the pattern across the individual states in Figure 4. Figure 6 shows the underlying data on average per capita income for the four regions. This figure shows clearly that the average incomes in each region have gotten much closer over time.[12] The main inference from Figures 5 and 6 is that a lot of the long–term reduction in σ_t shown in Figure 4 reflects the typical southern and western state becoming more like the typical eastern and midwestern state.

Figure 7 shows the patterns for σ_t within each of the four regions. The long–term decline in σ_t among the western states is apparent, but the other patterns are less straightforward. One clear result, however, is that the values of σ_t within each of the four regions are essentially the same toward the end of the sample.

The regional patterns highlight the distinction between σ and β convergence that we discussed before. With respect to σ convergence, the narrowing of the gap in average incomes across regions is a central element of the story and the changes within regions are a side show. In contrast, the estimated speeds of β convergence between and within regions are virtually identical. That is, relatively poor eastern states (such as Maine and Vermont in 1880) tended to catch up to relatively rich eastern states (Massachusetts and Rhode Island in 1880) about as fast as the typically poor southern state tended to catch up to the typically better off western or eastern state. These findings are consistent with the underlying model if the initially high values of σ_t reflected temporary disturbances that affected entire regions (such as the Civil War and the opening up of territories in the west).

Net Migration across U.S. States

This section, which extends the work of Sala–i–Martin (1990, Ch. 5), examines the migration of persons among the U.S. states in the context of the process of growth and convergence that we have been considering. As already mentioned, the process of convergence is quickened by movements of people from areas where ratios of capital to workers—and hence, wage rates and levels of per capita income—are low to areas where they are high. We investigate whether the flows of net migration accord with this story and whether these flows appear to be a substantial element in the extent of the convergence that we have estimated for the U.S. states.

Suppose that people (and hence, workers) are identical and that states offer the same amenities, government policies, and so on. Suppose that places differ by initial ratios of physical capital to labor and hence, wage rates, and that existing capital cannot move. Then people are motivated to move from low–wage to high–wage areas.

If moving were costless, then the migration of persons would equalize per capita incomes instantaneously in this model. In fact, moving entails costs, which include direct outlays for transportation, costs of familiarizing oneself with new jobs and surroundings, psychic costs of leaving acquaintances, and so on. If we allow for heterogeneity among persons, then the costs of moving differ in accordance with age, family status, occupation, and other characteristics that affect direct costs of moving or preferences about moving. Therefore, not all persons in low–wage areas are motivated to leave at a given point in time (or ever). This conclusion is reinforced if we allow for heterogeneity of jobs and workers so that wage rates and employment involve features of a matching problem.

The costs of moving into an area may depend on the aggregate flow of persons into that area. This rate of flow could influence job–search costs (which would show up in properly defined wage rates of new entrants) and also costs of housing. Thus, even if we abstract from matching considerations, these elements imply that not all migrants will go to the same place at a point in time.

If places differ in amenities that affect utility or production, such as climate, natural resources, and (exogenous?) government policies, then the long–run equilibrium described by Roback (1982) entails a distribution of wage rates for identical workers, along with a distribution of population densities and land prices. Although wage rates differ across places, these variations compensate for differences in land prices and amenities and people have no incentive to move. In terms of the reduced form, the equilibrium wage rate and population density for state i, w_i^* and π_i^*, are determined along with the land prices by the underlying amenities, denoted θ_i. We can think, as an approximation, of people having an incentive to move to state i if $\pi_i < \pi_i^*(\theta_i)$, so that $w_i > w_i^*(\theta_i)$. With costs of moving, the rate of migration into state i would be a positive function of the gap, $w_i - w_i^*(\theta_i)$, and the derivative of this function would be finite.

The analysis is more complicated with variable capital stocks. We are especially interested in analyzing the effects in the context of our data set, which includes information about per capita incomes (which we take as proxies for wage rates) and population densities, but not about capital stocks. If starred variables denote steady-state values, then a place with a temporarily high intensity for physical capital could have $w_i > w_i^*(\theta_i)$ and $\pi_i > \pi_i^*(\theta_i)$. For given $w_i - w_i^*(\theta_i)$, a higher $\pi_i - \pi_i^*(\theta_i)$ signals that current capital intensity is higher and hence, that capital intensity and wages rates will decline over time. In particular, the greater $\pi_i - \pi_i^*(\theta_i)$ the shorter the expected persistence of the gap between w_i and $w_i^*(\theta_i)$ and hence, the lower the incentive to migrate into the state.

The above reasoning leads us to write a function for m_{it}, the net rate of migration into state i between years t–T and t, as

$$(3) \qquad m_{it} = f(y_{i,t-T}, \theta_i, \pi_{i,t-T}; \text{ variables that depend on t but not i})$$

where the partial effects of $y_{i,t-T}$ and θ_i are positive (if a higher θ_i means more amenities) and the partial effect of $\pi_{i,t-T}$ is negative.[13] We assume that θ_i—an exogenous characteristic like climate or geography—does not change over time. (Thus, the analysis would have to be modified for exhaustible resources like silver or oil that get depleted over time or for changing government policies.) The set of variables that depend on t but not i includes any elements that influence the national averages of per capita income and population density; that is, $y_{i,t-T}$ and $\pi_{i,t-T}$ in equation (3) involve comparisons with alternatives available in other locations at time t. The set also includes effects like technological progress in heating and air conditioning that affect people's attitudes about climate or other components of the amenities, θ_i.

We have found empirically that a simple functional form of equation (3) does reasonably well in explaining net migration across states:

$$(4) \quad m_{it} = a + b \cdot \log(y_{i,t-T}) + c_1 \cdot \theta_i + c_2 \cdot \pi_{i,t-T} + c_3 \cdot (\pi_{i,t-T})^2 + v_{it}$$

where v_{it} is an error term, $b>0$, and the form allows for a quadratic in population density, $\pi_{i,t-T}$.[14] The marginal effect of $\pi_{i,t-T}$ on m_{it} is negative if $c_2+2c_3<0$. Although there is an extensive literature about variables to include in θ_i,[15] the present analysis includes only the log of average heating-degree days, denoted $\log(HEAT_i)$, which is a disamenity so that $c_1<0$. The variable $\log(HEAT_i)$ has a good deal of explanatory power for net migration; we did explore different functional forms and the addition of cooling–degree days as an explanatory variable.[16] But the alternative functional forms do not fit as well as the one described in equation (4) and the cooling–degree days variable is insignificant. Other components of the vector θ_i would be important for a fuller study of migration. It would also be useful to introduce migration for retirement, a mechanism that likely explains some outliers like Florida. However, these kinds of modifications probably would not change the basic findings that we now present about the relation between net migration and state per capita income and the interaction between migration and the convergence results.

Figure 8 shows the simple, long–term relation between in–migration and initial per capita income. The variable on the vertical axis is the average annual migration rate for each state from 1900 to 1987.[17] The horizontal axis plots the log of state per capita personal income in 1900. The figure shows a positive relationship (correlation = 0.51), but the relation is not nearly as clear–cut as that seen before for long–term per capita growth in Figure 1. Figure 9 shows the partial relation between the long–term migration rate and the log of initial per capita income (after holding constant the 1900 values of the right–side variables contained in equation (4), as well as regional dummies and the agricultural–share variable for 1900). The partial correlation is positive and equal to 0.45.

Table 2 shows regression results in the form of equation (4) for net migration into U.S. states.[18] The results are for eight sub–periods, beginning with 1900–1920.[19] The dependent variable is the ratio of migrants (annual average over each sub–period) to state population at the start of the sub–period.[20] Hence, the dependent variable approximates the contribution of net migration to the state's growth rate of population over the sub–period.

The equations include period–specific coefficients for $\log(y_{i,t-T})$ and $\log(HEAT_i)$, but single coefficients for the two population–density variables, $\pi_{i,t-T}$ (thousands of persons per square mile of total area) and $(\pi_{i,t-T})^2$. The regressions also include period–specific coefficients for regional dummies, agriculture share in personal income, $AGRY_{i,t-T}$, and (for sub–periods that start in 1930 or later) the variable structure, S_{it}. These variables are the ones used before for income growth in column 3 of Table 1. (The estimated coefficients of these other variables—not shown in Table 2—are sometimes statistically significant but play a relatively minor role overall.) The hypothesis of coefficient stability for the population-density variables is accepted at the 5% level (p–value = 0.34) and the other results change little if period-specific coefficients on these variables are introduced. The hypothesis of stability over the sub–periods in the coefficients of $\log(HEAT_i)$ is rejected at the 5% level (p–value = 0.000), although the estimated coefficients, $\hat{b}$, on $\log(y_{i,t-T})$ change little if only a single coefficient is estimated for $\log(HEAT_i)$.

The estimated coefficients of $\log(HEAT_i)$ in Table 2 are all negative and most are significantly different from zero. These results indicate that, *cet. par.*, people prefer warmer states. For population density, $\pi_{i,t-T}$, the jointly estimated linear term is significantly negative, –0.0452 (0.0077), and the square term is significantly positive, 0.0340 (0.0092). These point estimates imply that, *cet. par.*, the marginal effect of population density on in–migration is negative except for a few observations with the highest densities (New Jersey and Rhode Island since 1960 and Massachusetts since 1970). Since the implied marginal effect of population density for these outliers is small and since we are fitting a quadratic approximation, the true effect of population density could be negative throughout.

Figures 8 and 9 showed that the long–term relation between migration rates and initial income is positive, but not very strong. The regression results, which are conditioned on the values of $\log(y_{i,t-T})$ and $\pi_{i,t-T}$ at the beginning of each sub–period, are considerably clearer. The estimated coefficient, $\hat{b}$, on $\log(y_{i,t-T})$ is significantly positive for every sub-period shown in Table 2. The joint estimate, $\hat{b}$, for the eight sub–periods is 0.0261 (0.0023), which implies a t–value over 11. Thus, the regressions provide strong statistical evidence that, *cet. par.*, higher per capita income leads to a greater rate of net in–migration. The estimates in Table 2 do, however, reject at the 5% level the hypothesis of stability in the b coefficients across the sub–periods (p–value = 0.02).

Recall that we do not use individual state deflators for personal income. We can, however, interpret the population-density variable as a proxy for housing costs; these differences are a major source of variation in the cost–of–living across states. Thus, the estimated coefficients of per capita personal income in Table 2 likely represent effects for given costs of housing. It turns out, however, that the jointly estimated coefficient of $\log(y_{t-T})$ is significantly positive (with a t–value of 9) even if the only other regressors in the equations are period–specific constant terms. Thus, the results suggest that the measured differences in nominal per capita personal income across states do reflect variations in *real* per capita income.

Although the relation between the rate of in–migration and lagged per capita income is positive and highly statistically significant (holding fixed our measure of amenities, population density, and some other variables), the magnitude of the relation is small. For example, the joint estimate for b implies that, *cet. par.*, an increase in a state's per capita personal income by 10% raises net in–migration only by enough to raise the state's population growth rate by 0.26 percentage points per year. The slow adjustment through net migration means (unless there is a substantial response of a state's fertility or mortality) that population densities do not adjust rapidly to differences in per capita income adjusted for amenities. Our previous results suggest that differences in per capita income tend themselves to be eliminated over time, but only at a rate of about 2% per year. Thus, disparities in per capita income also persist for a long time. Putting these results together, the implication is that net–migration rates would be highly persistent over time. The data accord with this conclusion. For example, the correlation of the average net migration rate from 1900 to 1940 with that from 1940 to 1987 is 0.70. Figure 10 depicts this pattern of persistence.

As discussed before, the migration of raw labor from poor to rich states speeds up the convergence of per capita income. That is, the estimated coefficients, $\hat{\beta}$, shown in Table 1, should include the impact of migration. We use for comparative purposes the joint estimate, $\hat{\beta} = 0.0210$, that applies to column 3 of Table 1 for the eight sub–periods used in Table 2. We can use the estimated response of migration to $\log(y_{i,t-T})$—$\hat{b} = 0.026$ from Table 2—to quantify the effect of migration on the convergence coefficient. We have modified the neoclassical growth model to include endogenous migration as a source of linkage between population growth and the log of per capita income. (We neglect here any endogeneity of fertility or mortality.) The effect of the migration channel on the convergence speed, β, depends in the model on the underlying parameters of preferences and technology and on the quantity of human capital that migrants possess. We use parameter values that are consistent with the estimated values of β and b and with information from other studies (see B/X). If we assume unrealistically that migrants have zero human capital, then—depending on the specification of the underlying parameters—we calculate that β without migration would have been in the interval between 0.014 and 0.016, instead of the estimated value, 0.021. Thus, migration can account for as much as a third of the estimated speed of convergence if we neglect the human capital of migrants. The role of migration is, however, considerably less if we allow for migrants' human capital. For example, if the typical migrant's human capital per person is half the total capital stock per person that prevails in the recipient state, then the computed β without migration is between 0.018 and 0.019. Hence, if we allow for a reasonable amount of human capital, then migration cannot explain much more than 10% of the estimated speed of convergence.

We now attempt to get a direct estimate of the effect of migration on convergence speed by entering migration rates into the growth–rate regressions. The expectation is that (exogenous) in–migration will have a negative effect on the per capita growth rate and that the addition of the migration rate as a regressor will lower the estimated β coefficient. We first enter the contemporaneous migration rate, m_{it}, into regressions of the type presented in column 3 of Table 1. We drop the first sub–period (1880–1900) and now consider only the eight sub–periods that begin in 1900. If we restrict the coefficient on m_{it} to be the same for the eight sub–periods, then the estimated coefficient on m_{it} is ***positive*** and significant: 0.098 (0.029). The joint estimate of β—0.0250 (0.0027)—is actually somewhat higher than the value, 0.0210, that arises when the migration rate is excluded from the regression. Thus, contrary to expectations, the estimated β convergence does not diminish if we hold fixed net migration rates. If we allow for separate coefficients on m_{it} for each sub–period, then all eight point estimates are positive; the hypothesis of coefficient stability over the sub–periods is accepted at the 5% level (p–value = 0.37). In any event, the joint estimate of β—0.0256 (0.0027)—is about the same as that with a single coefficient for m_{it}.

A state's per capita growth rate and net migration rate are simultaneously determined. Suppose, for example, that a state is known to have favorable prospects for growth, but that these prospects are not adequately captured by the explanatory variables that we have included in the regressions for growth and migration. Then the residuals in each equation would tend to be positive; the positive residual in the migration equation reflects the response of migrants to the favorable growth opportunities that we do not hold constant with the included regressors. It seems likely that the positive estimated coefficients for m_{it} in the growth–rate regressions reflect this type of interaction.

We have also estimated by instrumental variables the growth–rate equations that include m_{it} as an explanatory variable. Aside from the (predetermined) variables that enter into the growth–rate equations in Table 1, column 3, we include as instruments the additional variables that influence the net migration rate in Table 2: $\log(HEAT_i)$, $n_{i,t-T}$, and $(n_{i,t-T})^2$.[21] If the coefficients on m_{it} in the growth–rate equations are restricted to be the same over the eight sub–periods, then the estimated coefficient of m_{it} is now 0.010 (0.047), which differs insignificantly from zero. The joint estimate of β—0.0214 (0.0030)—is close to the value, 0.0210, found when the migration rate is omitted from the regression. The findings are basically the same if we allow for separate coefficients on m_{it} for each sub–period. In particular, the joint estimate of β is 0.0209 (0.0032). These results suggest that exogenous shifts in net migration rates do not have a strong contemporaneous interaction with per capita growth: if we hold fixed exogenous net migration rates, then we estimate about the same rate of β convergence as we did before. These results should, however, be compared with the values of β that we expect to find when we hold constant the migration rates: between 0.014 and 0.016 if migrants have zero human capital and between 0.018 and 0.019 if human capital per person for migrants is half the total capital stock per person for the prior residents. The differences between the estimated coefficient, 0.0209 (.0032), and the predicted coefficients are small (and statistically insignificant) for the values that allow for human capital. Thus, the results are consistent with the modified neoclassical growth model that includes endogenous migration.

To summarize the main points on migration, we find that, *cet. par.*, per capita income has a highly significant positive effect on net migration rates into a state. Thus, we verify the predicted response of net migration to economic opportunities. We find, however, little contemporaneous interplay between net migration and economic growth. Specifically, we observe little change in estimated β coefficients when we hold constant net migration rates. These results are consistent with a modified neoclassical growth model that allows for endogenous migration; in particular, given the estimated response of migration to per capita income, the modified model predicts that migration would explain only a small part of β convergence.

Gross State Product

Data on gross state product (GSP) for 48 states are available from 1963 to 1986.[22] GSP, analogous to gross domestic product, refers to the payments to the factors that produce goods within a state, whereas personal income pertains to the returns to the factor owners, who may reside in other states. The main distinction between GSP and personal income arises for income from physical capital.

Table 3 shows regressions for per capita GSP for four sub–periods: 1963–1969, 1969–1975, 1975–1981, and 1981–1986. The concept of GSP in this table is the nominal aggregate for the state divided by the national deflator for GSP.[23] These figures reflect the current returns to factors of production and are therefore relevant for decisions on investment, migration, and so on. However, the measured growth rates pick up a combination of changes in quantities produced and changes in relative prices across sectors. The effects of the relative–price changes, which interact with the composition of production within a state, can be viewed as part of the error term, u_{it} in equation (1), that is filtered out by the structural–composition variable, S_{it}, that we discussed before. For GSP, the variable S_{it} is based on a division of production into 54 sectors. Thus, the breakdown is much finer than the 9–sector construct used for personal income in Table 1.

Overall, the results on β convergence for GSP from Table 3 are similar to those for personal income from Table 1. If we exclude the structure variable and include only lagged GSP (column 1 of Table 3) or if we add regional dummies (column 2), then the estimated coefficients $\hat{\beta}$ are unstable. The estimates are far more stable when we add the explanatory variable S_{it} in column 3. The joint estimate of β for the four sub–periods is 0.0216 (0.0042) and the hypothesis of stability in the β coefficients over the four sub–periods is accepted at the 5% level (p–value = 0.64).

Figure 11 shows a plot of the average growth rate of per capita GSP from 1963 to 1986 against the log of per capita GSP in 1963. The downward–sloping relation is clear, although the fit is not as good as that for the long period relation for personal income shown in Figure 1. The main difference relates to the sample period and not to the distinction between GSP and personal income.

Table 4, which extends the analysis of Sala–i–Martin (1990, Ch. 3), breaks down the results by sectors of gross state product. We look at GSP per worker originating in eight standard non–agricultural sectors: mining, construction, manufacturing, transportation, wholesale & retail trade, finance–insurance–real estate (FIRE), services, and government.[24] We have omitted the agriculture sector because data on agricultural employment are not comparable to those for the non–agricultural sectors. The first two columns of the table show the shares of each sector in U.S. aggregate GSP in 1963 and 1986. The main patterns in the shares, which are well known, are the declines in manufacturing and agriculture (the residual from the sum of the eight sectors) and the increases in services and FIRE.

Table 4 shows positive estimates $\hat{\beta}$ for each of the eight sectors over the period 1963–1986, although not all of the estimates are statistically significant. (Each of these regressions includes a constant, the log of the sector's productivity in 1963, and the regional dummies.) Basically, the $\hat{\beta}$ values for the four service–type sectors—trade, FIRE, services, and government—are similar and fall in a range from 0.009 to 0.016. The $\hat{\beta}$ values are higher for the other four sectors, especially for manufacturing where the estimate is 0.0460 (0.0082). It is only this high and precisely estimated value for manufacturing that leads to rejection of the hypothesis that the β coefficients are the same across the sectors. The joint estimate of β for the eight sectors is 0.0213 (0.0024), but we reject at the 5% level the hypothesis that the individual β's are the same (p–value = 0.002). We would accept the hypothesis that the β coefficients are the same for the seven sectors other than manufacturing—the estimate is $\hat{\beta}$=0.0164 (0.0024) and the p–value for the test of equality for the coefficients is 0.77.

The main inference that we draw from Table 4 is that β convergence applies within sectors in a manner that is broadly similar to that found in Tables 1 and 3 for state aggregates of personal income and gross state product. Thus, an important part of the overall process of convergence across the states involves adjustments of productivity levels within sectors.

Figure 12 shows the (unweighted) cross–sectional standard deviation, σ_t, for the log of per capita GSP from 1963 to 1986. The decline of σ_t from 0.18 in 1963 to a low point of 0.13 in 1972 accords with the behavior for personal income shown in Figure 4. We think that the rise in σ_t for GSP to a peak of 0.18 in 1981 reflects the behavior of oil prices. Especially for 1979–1981, the oil shocks benefit the states that already have above–average GSP per capita and thereby lead to an increase in σ_t. After 1981, the decline in σ_t reflects the normal pattern of σ convergence, reinforced later by a fall in oil prices.

The different patterns from 1973 to 1986 in σ_t based on GSP versus σ_t based on personal income reflect, at least in part, differences in the relation between shares of product or income originating in oil–related industries and the levels of per capita product or income. The correlation of the log of per capita GSP with the share of GSP originating in crude oil and natural gas rises because of the oil shocks from 0.1 in 1973 to 0.4 in 1975 and 0.7 in 1981, and then falls with the decline in oil prices to 0.1 in 1986. In contrast, the correlation of the log of per capita personal income with the share of personal income originating in oil and natural gas is –0.3 in 1970 and 0.0 in 1980. These divergent patterns reflect the distinction between the location of oil and gas facilities and the ownership of these facilities.

From 1973 to 1981 the oil shocks have less of an effect on σ_t for personal income than for GSP because, unlike for GSP, the rises in oil prices do not particularly harm the states with already low levels of per capita personal income. Similarly, a possible reason why σ_t for personal income does not decline later in the 1980s is that, unlike for GSP, the declines in oil prices do not particularly benefit the states with low per capita incomes.

Regions of Europe

We now apply the analysis to the behavior of gross domestic product in regions of seven countries in Europe. We have data on GDP and a few other variables for 73 regions: 11 in Germany, 11 in the United Kingdom, 20 in Italy, 21 in France,[25] 4 in the Netherlands, 3 in Belgium, and 3 in Denmark. Table 5 shows the breakdown of the regions.

Data for 1950, 1960, and 1970 are from Molle (1980).[26] Data for 1966 (missing France and Denmark), 1970 (missing Denmark), 1974, 1980, and 1985 are from Eurostat (various issues). The nominal figures on GDP are expressed via current exchange rates in terms of a common currency unit. It is unnecessary to deflate the nominal values for the purposes of the cross–section equations that we consider; that is, any common deflation affects only the constant terms in the regressions.[27] Aside from GDP and population, the data set includes a breakdown of employment into three sectors—agriculture, industry, and services—for 1950, 1960, and 1970, and a breakdown of GDP into the same three sectors for 1966 (missing France and Denmark), 1970, 1974, 1980, and 1985. (The data for Denmark on the breakdown of GDP are available only for 1974.)

Figure 13 shows for the 73 regions the relation of the growth rate of per capita GDP from 1950 to 1985 to the log of per capita GDP in 1950. (The numbers of the regions correspond to those in Table 5. See Figure 14 for a map that shows the locations of the regions.) The values shown are all measured relative to the means of the respective countries. The figure shows the type of negative relation that is familiar from the study of the U.S. states. The correlation between the growth rate and the log of initial per capita GDP in Figure 13 is –0.70.

Because the underlying numbers are expressed relative to own–country means, the relation in Figure 13 pertains to β convergence within countries rather than between countries. For the seven countries that we are considering, the estimates of β convergence between countries turn out to be similar to those within the countries. Previous research, which includes Baumol (1986), DeLong (1988), Dowrick and Nguyen (1989), and Barro (1991), has considered β and σ convergence among larger groups of countries. Since the seven–country data set considered here provides much less information about behavior across countries, we shall focus our attention on the within–country results.

Table 6 shows regressions for the European regions over four sub–periods: 1950–1960, 1960–1970, 1970–1980, and 1980–1985.[28] The form of the analysis parallels that for the U.S. states from Tables 1 and 3. The regressions in column 1 of Table 6 include only a constant and $\log(y_{i,t-T})$ as independent variables. The estimated coefficients, $\hat{\beta}$, are positive but unstable across the periods. The pattern of results over the sub–periods is similar to that found for the U.S. states, and the joint estimate, 0.0183 (0.0029), is slightly smaller than that found before. The hypothesis of a constant β coefficient is again rejected at the 5% level (p–value = 0.000).

Column 2 of Table 6 adds country dummies, which have enormous explanatory power for growth rates of European regions. We think of the country dummies, analogous to the regional dummies that we used for the United States, as proxies in equation (1) for the steady–state values, x_i and $\hat{y}_i^*$, and for countrywide fixed effects in the error term, u_{it}. The addition of the country dummies in column 2 makes the $\hat{\beta}$ coefficients markedly more stable across the sub–periods, but the joint estimate—0.0186 (0.0021)—is very close to that shown in column 1. (This joint estimation includes period–specific country dummies.) The results in column 2 still reject at the 5% level the hypothesis of equal β coefficients across the sub–periods (p–value = 0.008).

The results with country dummies in column 2 show within–country β convergence and are analogous to that shown in Figure 13. In contrast, the results from column 1 show a combination of within– and between–country β convergence. The joint estimates $\hat{\beta}$ in the two columns are similar because the rates of within– and between–country β convergence are nearly the same in this seven–country sample. We can also estimate β by using only the data on country aggregates (just as we did for the U.S. regions). Then the jointly estimated β coefficient for the four sub–periods shown in Table 6 turns out to be 0.0183 (0.0029), virtually the same as the value, 0.0186 (0.0021), shown in column 2. Note that the first value, 0.0183, is an estimate of β convergence between countries, whereas the second, 0.0186, is an estimate within countries.

Column 3 of Table 6 adds the shares of agriculture and industry in total employment at the start of the sub–period for the 1950–1960, 1960–1970, and 1970–1980 sub–periods. The regression for the 1980–1985 sub–period adds the shares in overall GDP at the start of the period. These share variables are analogous to the agricultural share and structural composition variables that we used before for the United States. In effect, the share variables for the European regions are as close as we can come with our present data to the structural variable, S_{it}, that we used for the United States.

The main new result from column 3 of Table 3 is the acceptance of the hypothesis of stability in the β coefficients at the 5% level (p–value = 0.46). (These results allow for period–specific coefficients on the share variables and the country dummies.) The joint estimate, $\hat{\beta} = 0.0178$ (0.0034), does not change much from that shown in column 2. This point estimate—showing β convergence at slightly below 2% per year—is somewhat less than the corresponding value, 0.0216, found for the U.S. states in Table 3.

We have also estimated the joint system with individual β coefficients for the seven countries. This system corresponds to the 4–period regression shown in column 3 of Table 6 except that the coefficient β is allowed to vary over the countries (but not over the sub–periods). Thus, the system contains period–specific country dummies and the agricultural and industrial share variables (with coefficients that vary over the sub–periods but not across the countries). The resulting estimates for β are:

Germany (11 regions):	0.0230 (0.0061)
United Kingdom (11 regions):	0.0337 (0.0093)
Italy (20 regions):	0.0118 (0.0036)
France (21 regions):	0.0097 (0.0059)
Netherlands (4 regions):	0.0496 (0.0202)
Belgium (3 regions):	0.0237 (0.0164)
Denmark (3 regions):	0.0018 (0.0211)

The likelihood–ratio statistic for equality of the β coefficients across the seven countries is 12.6, which coincides with the 5% critical value from the χ^2 distribution with 6 df. We could try to come up with reasons why the regions in the Netherlands and the United Kingdom have higher than average β convergence, whereas those in Denmark, France, and Italy have lower than average convergence. But, since the differences are only marginally significant in a statistical sense, the main conclusion is that similar rates of β convergence are consistent with the data.

Figure 15 shows the (unweighted) standard deviation, σ_t, for the log of per capita GDP (expressed relative to the mean for the respective country) for the 73 European regions. (The data point for 1966 is based on partial coverage because figures for France and Denmark are unavailable.) Since we filtered out the country means, the values shown in the figure refer to σ convergence for regions within countries and not across countries. The principal observation is that σ_t for the European regions declined from 0.28 in 1950 to 0.18 in 1985. The value for Europe in 1985 is still somewhat above the value 0.14 for U.S. GSP in 1986 (or the low point of 0.13 in 1972).

Figure 15 shows that the fall in σ_t for the European regions moderated from 1974 to 1985. We found somewhat similar behavior for σ_t based on U.S. GSP in Figure 12, although the U.S. results showed a substantial rise in σ_t from the mid 1970s until the early 1980s. For the United States, we think that we can explain part of the pattern in σ_t after the mid 1970s from the behavior of oil shocks; a similar story may account for the behavior of σ_t for Europe in Figure 15. (Although the United Kingdom is the only oil producer among the seven countries, the regions of Europe can still vary substantially in their sensitivity to oil shocks.)

Figure 16 shows the behavior of σ_t for the regions within the four largest European countries in the sample: Germany, the United Kingdom, Italy, and France. The countries are always ranked, highest to lowest, as Italy, Germany, France, and the

United Kingdom. The overall pattern shows declines in σ_t over time for each country, although little net change occurs since 1970 for Germany and the United Kingdom. In particular, the rise in σ_t from 1974 to 1980 for the United Kingdom—the one oil producer in the European sample—likely reflects the effects of oil shocks. In 1985, the values of σ_t are 0.26 for Italy, 0.20 for Germany, 0.15 for France, and 0.10 for the United Kingdom, compared with 0.15 for U.S. GSP. Thus, although σ_t for Italy has fallen from 0.42 in 1950, Italy still has a way to go to attain the regional dispersion of per capita GDP that is characteristic of the other countries.

The high value of σ_t for Italy reflects especially the spread between the prosperous north and the poor south. A popular view, in fact, is that the backward regions of southern Italy will always lag behind the advanced regions of northern Italy (and vice versa for the United Kingdom). Our overall findings do not accord with this type of story since we find substantial evidence of β and σ convergence across the regions of Europe. For example, with respect to β convergence in Figure 13, many of the observations with the highest initial per capita GDP (relative to the own–country mean) are for northern Italy, whereas many with the lowest per capita GDP are for southern Italy (see Table 5 and the map in Figure 14). These observations scatter reasonably well around the regression line; that is, as predicted, the initially poorer regions in Italy grow faster on average than the initially richer.

Table 7 shows the full array of data for averages of four prosperous regions in northern Italy and seven poor regions in southern Italy. The northern regions began in 1950 with per capita GDP 70% above the mean for Italy, whereas the southern regions began 47% below the mean. As predicted, the northern regions grew from 1950 to 1985 at a below–average rate—0.71% per year below the mean—whereas the southern regions grew at an above–average rate—0.39% per year above the mean. Accordingly, in 1985, the northern regions were only 38% above the mean, whereas the southern regions were only 34% below the mean. The relative growth performances from 1950 to 1985 correspond well to the predicted behavior implied by the jointly estimated value, $\hat{\beta}$=0.0178 per year, from column 3 of Table 6. That value implies that the northern regions should have grown on average at a rate 0.70% per year below the mean, whereas the southern regions should have grown on average at a rate 0.51% per year above the mean. Thus, there is no indication of something out of the ordinary in the relative performances of the regions of north and south Italy. The south of Italy has not caught up to the north because it started far away and the rate of β convergence is only about 2% per year.

Table 7 shows comparable statistics for the north and south of Great Britain. (The region for Northern Ireland—a substantial outlier for the United Kingdom—is excluded in these calculations.) One immediate observation is that the magnitude of the dispersion between the south and north of Great Britain is trivial relative to that between the north and south of Italy. In any event, because the six northern and four southern regions in Great Britain began in 1950 with similar averages for per capita GDP, the theory predicts that subsequent growth rates would also be similar. In fact, the north grew by 0.05% per year below the mean, whereas the south grew at 0.07% per year above the mean. Therefore, in 1985, the average level of per capita GDP in the north was about 3% below the mean, whereas that in the south was about 5% above the mean. The theory does not predict this outcome, which can likely be explained by sectoral disturbances that affected the regions differentially (and in a way that was uncorrelated with the initial levels of per capita GDP).

The Results on β Convergence

A striking aspect of our findings is the similarity in the estimated rates of β convergence in different contexts. We first summarize the elements of this empirical regularity, then assess the similarity in the estimates from a theoretical perspective, and finally show the significance of the results by applying them to developments in recently unified Germany.

We find ample evidence that poorer regions within a country tend to grow faster than richer regions, a property that we call β convergence. For U.S. per capita personal income from 1880 to 1988, we estimate β to be around 2% per year whether we look within or across four major geographical regions. We also get similar estimates of β when we examine per capita gross state product (GSP) from 1963 to 1986. For GSP, β convergence appears within eight standard non-agricultural sectors of production (mining, construction, manufacturing, transportation, wholesale & retail trade, fire-insurance-real estate, services, and government), although the size of β for manufacturing is substantially higher than those for the other sectors.

The results for 73 regions of 7 European countries (Germany, the United Kingdom, Italy, France, the Netherlands, Belgium, and Denmark) apply to per capita GDP from 1950 to 1985. The estimated rates of β convergence are similar to those found for the United States; in particular, we see no evidence that poor regions, such as those in southern Italy, are being systematically left behind in the growth process. For the seven countries considered in this study, the cross–country estimates of β are similar to the within–country estimates.

We have obtained estimates of β for a broader cross section of countries in the post–World War II period: one sample contains 20 OECD countries and another comprises a less homogeneous group of 98 countries. If we examine only the simple relation between the per capita growth rate and initial per capita GDP, then the estimates of β are around 1% per year for the OECD sample and about 0 for the larger sample. Recall, however, that the neoclassical growth model summarized by equation (1) predicts a conditional form of convergence in which differences in per capita product enter relative to differences in steady–state positions, $\hat{y}_i^*$ and x_i. If we hold constant additional variables that we interpret as proxies for differences across countries in steady–state positions, then we again obtain estimates of β in the neighborhood of 2% per year (see Barro and Sala–i–Martin [1991, Table 5]). These results suggest that the ranking of the divergence in the steady–state values, $\hat{y}_i^*$ and x_i, goes from the heterogeneous collection of 98 countries at the top to the relatively homogeneous OECD countries to the still more homogeneous regions within the United States or the seven European countries. In the regional context, our long–period estimates of β depend little on whether we hold constant proxies for steady–state values, a result that suggests little regional variation of steady–state values within the countries that we have studied.

The neoclassical growth model does not imply that the convergence coefficient, β, would be the same in all times and places. The coefficient depends, as we discussed before, on the underlying parameters of technology and preferences, but not on differences in technologies or government policies that can be represented as proportional effects on the production function, that is, as variations in the parameter A in the function, $Af(\hat{k})$. These A–type effects have important influences on steady–state output per worker, $\hat{y}_i^*$, but not on the speed with which an economy approaches its steady state. Therefore, economies that differ greatly in some respects may nevertheless exhibit similar rates of β convergence.

We noted that a greater degree of labor mobility leads theoretically to a higher convergence coefficient β. This effect means that the rates of β convergence would be higher for regions of countries than across countries. Direct estimates for the effect of net migration across the U.S. states indicate, however, that this effect is small. In particular, the magnitude of the effect is not large enough to generate a statistically detectable gap between the β coefficients for regions and countries.

Capital mobility also tends to be greater across regions than across countries. The effects of capital mobility on β convergence are, however, difficult to pin down. With identical technologies, capital mobility speeds up convergence for per capita product but slows down convergence for per capita income. Our results for the U.S. states show little distinction in the dynamics of product and income, an observation that induces us to deemphasize capital mobility. Also, if technologies (including government policies) differ across economies, then capital may move from poor to rich economies and lead thereby to divergence of per capita product. Thus, it is not obvious that greater capital mobility across regions than across countries would lead to higher rates of β convergence for regions than countries.

Suppose that, despite the theoretical ambiguities, we take it as an empirical regularity that the rate of β convergence is roughly 2% per year in a variety of circumstances. We can highlight the potential significance of this finding by showing how it applies to the recent merger of East and West Germany, the topic of the paper at this conference by Akerlof, et al (1991). Suppose that the ratio of the west's per capita income to the east's in 1990 is two, the order of magnitude suggested by Akerlof, et al. A coefficient β of 2% per year implies that the east's per capita income would grow initially by 1.4% per year higher than the west.[29] The half–life of this convergence process is 35 years; that is, it would take 35 years for half of the initial east–west gap to be eliminated. Thus, the results extrapolated from our findings for regions of the United States and Europe and for a variety of countries imply that anything close to "parity" in the short run is unimaginable.

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Footnotes

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[1]Some of the counterfactual results in open–economy models with perfect capital markets disappear if we assume that people become less patient as they raise assets and consumption (see Uzawa [1968]). This form of preferences is introspectively unappealing, but Blanchard (1985) shows that the aggregation across finite–lived individuals makes overall economies act this way. In particular, the initially most patient economy stops short of owning everything in the long run and the less patient economies do not tend toward zero consumption per effective worker. Assets are, however, still likely to become negative for less patient economies.

[2]Convergence can be less rapid if immigrants to rich economies are substantially above average in human capital. See Borjas (1990) for a discussion of the characteristics of immigrants.

[3]Quah (1990) discusses β and σ convergence in terms of Galton's Fallacy: the observation that heights of persons in a family regress to the mean across generations (a form of β convergence) does not imply that the dispersion of heights across the population diminishes over time (which would be an example of σ convergence). None of this makes β convergence uninteresting, as Quah seems to suggest; it just points out that β and σ

convergence are different concepts. One example of the Quah–Galton effect is the ordinal rankings of teams in a sports league. Although σ_t is constant by definition (so σ convergence cannot apply), we can still think of β convergence in terms of how rapidly teams at the bottom of the ranking tend to rebound toward the middle or how quickly champions tend to revert to mediocrity. The sports example also leads naturally to the issue of overshooting: is a currently weak team or country likely to be in a better position than a currently strong team or country at some future date? In the Quah–Galton context, if person 1 is taller than person 2, would we predict that the offsprings of person 2 would eventually be taller than those of person 1? This type of overshooting cannot obtain in the standard neoclassical growth model, which generates a first–order differential equation that is approximated in the linear, log–difference form of equation (1). Thus, if x_i and y_i^* are the same for all i, then if economy 1 starts out ahead of economy 2 we would predict that economy 1 would still be ahead of economy 2 at any date in the future. Our conjecture is that heights also satisfy this property, although we have not examined the data. The possibility of overshooting seems more likely for the rankings of sports teams; in fact, this area may be the best place to apply models of overshooting. For an overview of these models, see *The Economist* (1990).

[4]The data on personal income are from Bureau of Economic Analysis (1986), recent issues of *U.S. Survey of Current Business*, and Easterlin (1957, 1960). See B/X for a discussion. There are no data for Oklahoma for 1880 (which preceded the Oklahoma land rush) and we exclude Alaska, Hawaii, and Washington, D.C. throughout the analysis. We use nominal income figures deflated by the overall CPI. If the price level is the same for all states at each point in time, then we can just as well use nominal income figures for the cross–sectional analysis that we carry out. If prices differ across states at a point in time—that is, if there are departures from purchasing–power parity—then it would be preferable to use individual–state deflators. We think, however, that the available price

indexes across states do not improve on the assumption of a common price level. (The analysis requires only constant relative prices for growth rates, but equal levels of prices for levels of real income.)

[5]These results come from iterative, weighted, non–linear least squares, which allows for heteroscedasticity across the sub–periods but not for correlation of the error terms over the periods. We have also estimated systems that allow for this correlation, using seemingly-unrelated regression or SUR. In most cases, the results of hypothesis tests are similar. In some cases, however, we had difficulty with convergence of the estimates because of the interaction of the non–linearity in the model with the large number of parameters introduced by the SUR procedure. Probably it would be better to estimate parsimonious representations that allow for a restricted form of serial correlation in the errors, rather than an arbitrary pattern across the sub–periods.

[6]The ratio of the WPI for farm products to the CPI for all items fell at a rate of 3.5% per year from 1920 to 1930. Over that period, the average growth rate of real per capita farm income—nominal income divided by the CPI and farm population—was –2.7% per year. In contrast, the average growth rate of real per capita non–farm personal income was 0.8% per year. The data are from U.S. Department of Commerce (1973, 1975).

[7]It is well known that temporary measurement error in y_{it} can lead to an overestimate of the convergence coefficient β. In previous research we have taken several approaches to assessing the likely magnitude of this effect (see Barro [1991] and B/X). In one approach we related the growth rate, $(1/T)\cdot\log(y_{it}/y_{i,t-T})$, to income at a date prior to t–T, say to $\log(y_{i,t-T-T'})$. If measurement error does not persist over an interval greater than T' (which we took to be five or ten years), then with a plausible magnitude for β, the asymptotic bias in this form is in the direction opposite to that in the original form.

Because the empirical estimates of β from the two forms did not differ greatly we argued that the effects of measurment error were unlikely to be major.

[8]We do not have reliable data on agricultural employment but the data on farm population suggest that this productivity differential is large, at least in earlier years. We measure farm productivity as farm national income divided by farm population and non–farm productivity as non–farm national income divided by non–farm population. Using these concepts of productivity, the ratio of non–farm to farm productivity was 4.0 in 1889, 2.7 in 1899, 2.3 in 1909, 2.9 in 1920, 3.6 in 1930, 3.7 in 1940, 2.4 in 1960, 1.8 in 1970, 1.6 in 1980, and 1.5 in 1988. The data are from U.S. Department of Commerce (1973, 1975) and *Statistical Abstract*, 1990. One shortcoming of these measures of productivity is that they do not adjust for differences in family size between farm and non–farm populations.

[9]The ratio of the WPI for farm products to the CPI for all items grew at an average annual rate of 9.5% from 1940 to 1950. Over this period, the average growth rate of real per capita farm income (nominal income divided by the CPI and farm populaton) was 7.8% per year, compared to 2.9% per year for real per capita non–farm personal income. The data are from U.S. Department of Commerce (1975).

[10]To estimate σ_u we use the first–difference equation for σ_t^2, derived in B/X, $\sigma_t^2 = \sigma_u^2/(1-e^{-2\beta}) + [\sigma_0^2 - \sigma_u^2/(1-e^{-2\beta})]\cdot e^{-2\beta t}$. The steady–state variance, σ^2, equals the first term in this formula, $\sigma_u^2/(1-e^{-2\beta})$.

[11]We have the data on transfers only from the Commerce Department data that begin in 1929. Since the amounts for earlier years are small, the behavior of σ_t with and without transfers would be similar before 1929.

[12]See Easterlin (1957) and Borts and Stein (1964, Ch. 2) for related analyses of the regional dispersion of per capita personal income.

[13]See Mueser and Graves (1990) for a related model of migration.

[14]Population density is the ratio of state population to total area (land plus water). The data on area are from U.S. Department of Commerce, *Statistical Abstract*, 1990.

[15]See, for example, Blomquist, Berger, and Hoehn [1988]. Some of the variables that they consider, such as criminal activity, are, however, not exogenous in the same sense as climate or geography.

[16]The data on heating– and cooling–degree days refer to average temperatures from 1951 to 1980 and are from U.S. Department of Commerce, *Statistical Abstract*, 1990.

[17]The variable is the average of the rates for the subperiods, 1900–1920, 1920–1930, ..., 1970–1980, 1980–1987, weighted by the lengths of each interval. The rate for each sub–period is the annual average of net migration divided by state population at the start of the sub–period.

[18]The regressions use iterative, weighted least squares.

[19]The overall results do not change greatly if we add the sub–period, 1880–1900. This sub–period includes some enormous rates of in–migration that correspond to the opening up of new territories. Because our simple functional form does not fit well in this range we decided to exclude this sub–period from the present analysis.

[20]The data on migration are from U.S. Department of Commerce (1975). Recent figures are computed from data on population, births, and deaths from U.S. Department of Commerce, *Statistical Abstract*, various issues.

[21]The assumption here is that the instrumental variables, $\log(\mathrm{HEAT}_i)$ and $n_{i,t-T}$, do not enter directly as influences on per capita growth.

[22]The data are from Renshaw, Trott, and Friedenberg (1988). See B/X for a discussion.

[23]Individual state deflators are unavailable. Since we use a common deflator at each point in time, the particular deflator that we use affects only the constant term in the regressions. See n. 4.

[24]The data on employment by sector are from the Bureau of Labor Statistics.

[25]We lost one region for France because some of the data on Corse are combined with those for Provence–Alpes–Côte d'Azur.

[26]We appreciate the suggestion from Carol Heim to look at these data.

[27]Departures from purchasing–power parity across countries would not affect our main results, which filter out own–country effects. The growth rates for regions within countries involve the same kind of sensitivity to changes in relative prices that applied to GSP for the U.S. states.

[28]We have two alternative sources of GDP for 1970—Molle (1980) and Eurostat—and the two sources do not coincide. We computed the figures for 1960–1970 from Molle and those for 1970–1980 from Eurostat. Since the correlation between the two measures of the levels of per capita GDP in 1970 is 0.988, this discrepancy should not be important.

[29]We can also use the findings for the United States (Table 2) to estimate net migration from the east of Germany to the west. The resulting estimate (which allows for the differences in per capita income and population density, but not for differences in amenities) is that 1.2% or 203,000 of the eastern population would migrate over a year to the west. Akerlof, et al (1991, Table 9) show that the net out–migration from the east averaged 22,800 per month over the three months since the unification in July 1990. Although the actual flow of 274,000 at an annual rate exceeds our estimate of 203,000, the extrapolation of the U.S. experience to Germany does provide a reasonable order of magnitude.

Table 1: Regressions for Personal Income across U.S. States

	(1) No other var's.		(2) Regional dums.		(3) Regional dums., AGRY, structure	
Period	$\hat{\beta}$	R^2 $[\hat{\sigma}]$	$\hat{\beta}$	R^2 $[\hat{\sigma}]$	$\hat{\beta}$	R^2 $[\hat{\sigma}]$
1880-1900	.0101 (.0022)	.36 [.0068]	.0224 (.0040)	.62 [.0054]	.0268 (.0048)	.65 [.0053]
1900-1920	.0218 (.0032)	.62 [.0065]	.0209 (.0063)	.67 [.0062]	.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	.0431 (.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]	.0173 (.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]
9 periods, β restricted	.0175 (.0013)	--	.0189 (.0019)	--	.0224 (.0022)	--

Likelihood-ratio statistic for equal β's (.05 χ^2 value with 8 df = 15.5):

	(1)	(2)	(3)
	65.6	32.1	12.4
p-value:	(.000)	(.000)	(.134)

Note: 48 observations, except 47 for 1880-1900 in columns 1 and 2 (excluding Oklahoma) and 46 in column 3 (also excluding Wyoming, which lacks data on AGRY for 1880). The regressions use non-linear least squares and take the form, $(1/T)\cdot\log(y_{it}/y_{i,t-T}) = a - [\log(y_{i,t-T})]\cdot[1-\exp(-\beta T)]/T$ + other variables, where T is the length of the interval and y_{it} is per capita personal income for state i at time t. The other variables in column 2 are regional dummies for south, midwest, and west. The regression in column 3 adds the share of personal income originating in agriculture at the start of the period and the structural-composition variable (9 sectors) described in the text. The 9-period regression with a single value for β comes from iterative, weighted non-linear least squares. Standard errors are in parentheses, $\hat{\sigma}$ is the standard error of estimate.

Table 2: Regressions for Net Migration into U.S. States

Period	$\hat{b}$	log(HEAT)	Density	Dens. sq.	R^2 $[\hat{\sigma}]$
1900-1920	.0335 (.0075)	-.0066 (.0037)	-.0452 (.0077)	.0340 (.0092)	.70 [.0112]
1920-1930	.0363 (.0078)	-.0124 (.0027)	"	"	.61 [.0079]
1930-1940	.0191 (.0037)	-.0048 (.0014)	" "	" "	.71 [.0042]
1940-1950	.0262 (.0056)	-.0135 (.0022)	"	"	.83 [.0065]
1950-1960	.0439 (.0085)	-.0205 (.0031)	"	"	.76 [.0091]
1960-1970	.0436 (.0082)	-.0056 (.0025)	"	"	.70 [.0069]
1970-1980	.0240 (.0091)	-.0076 (.0024)	"	"	.73 [.0071]
1980-1987	.0177 (.0057)	-.0075 (.0018)	"	"	.73 [.0049]
8 periods, β restricted	.0261 (.0023)	indiv.	-.0447 (.0078)	.0329 (.0093)	--

Likelihood-ratio statistic for equal b's is 17.0 (.05 χ^2 value with 7 df = 14.1, p-value = .017)

Note: 48 observations. The regressions use iterative, weighted least squares and take the form: $RMG_{it} = a + b \cdot \log(y_{i,t-T}) + c_1 \cdot \log(HEAT_i) + c_2 \cdot DENS_{i,t-T} + c_3 \cdot (DENS_{i,t-T})^2$ + other variables. The coefficients c_2 and c_3 are constrained to be the same for all sub-periods. RMG_{it} is the average annual net migration into state i between years t-T and t, expressed as a ratio to the state's population in year t-T. $y_{i,t-T}$ is per capita personal income as in Table 1. $HEAT_i$ is average heating-degree days for state i (formed as an average for available cities in the state). $DENS_{it}$ is population density (thousands of persons per square mile of total area) for state i in year t. Other variables are the regional dummies, agriculture share in personal income, and sectoral-composition variables, as discussed in the notes to Table 1. The 8-period regression constrains the value of b to be the same for all sub-periods. Standard errors are in parentheses, $\hat{\sigma}$ is the standard error of estimate.

Table 3: Regressions for Gross State Product across U.S. States

	(1)		(2)		(3)	
	No other var's.		Regional dums.		Regional dums., structure	
Period	$\hat{\beta}$	R^2 $[\hat{\sigma}]$	$\hat{\beta}$	R^2 $[\hat{\sigma}]$	$\hat{\beta}$	R^2 $[\hat{\sigma}]$
1963-1969	.0317 (.0067)	.36 [.0070]	.0154 (.0060)	.63 [.0056]	.0157 (.0060)	.63 [.0056]
1969-1975	.0438 (.0166)	.16 [.0138]	.0406 (.0162)	.41 [.0120]	.0297 (.0101)	.74 [.0081]
1975-1981	-.0159 (.0133)	.03 [.0145]	-.0285 (.0134)	.17 [.0139]	.0258 (.0108)	.78 [.0072]
1981-1986	.1188 (.0294)	.39 [.0205]	.1130 (.0251)	.62 [.0168]	.0238 (.0091)	.92 [.0079]
4 periods, β restricted	.0335 (.0057)	--	.0211 (.0053)	--	.0216 (.0042)	--
Likelihood-ratio statistic for equal β's (.05 χ^2 value with 3 df = 7.8):	75.6		31.2		1.7	
p-value:	(.000)		(.000)		(.637)	

Note: 48 observations. The regressions use non-linear least squares and take the form: $(1/T)\cdot\log(y_{it}/y_{i,t-T}) = a - [\log(y_{i,t-T})]\cdot[1-\exp(-\beta T)]/T +$ other variables, where T is the length of the interval and y_{it} is per capita gross state product at time t. The other variables in column 2 are regional dummies for south, midwest, and west. The regression in column 3 adds the structural-composition variable (54 sectors) described in the text. The 4-period regression with a single value for β comes from iterative, weighted non-linear least squares. Standard errors are in parentheses, $\hat{\sigma}$ is the standard error of estimate.

Table 4: Regressions for Sectors of Gross State Product

	U.S. Share of Sector				
	1963	1986	$\hat{\beta}$	R^2	$\hat{\sigma}$
Mining	.023	.022	.0240 (.0074)	.49	.0134
Construction	.048	.047	.0169 (.0203)	.20	.0110
Manufacturing	.284	.199	.0460 (.0082)	.73	.0041
Transportation	.092	.094	.0257 (.0176)	.15	.0045
Wholesale & Retail Trade	.164	.169	.0093 (.0064)	.24	.0030
Finance, Insur., Real Estate	.145	.167	.0149 (.0077)	.43	.0046
Services	.105	.166	.0149 (.0077)	.27	.0036
Government	.102	.115	.0161 (.0039)	.55	.0032
8 Sectors Jointly	.963	.978	.0213 (.0024)	--	--

Likelihood-ratio Statistic for equal β's is 22.4 (.05 χ^2 value with 7df = 14.1, p-value = .002)

Note: 48 observations except 42 for mining. The regressions use non-linear least squares and take the form: $(1/T)\cdot\log(y_{it}/y_{i,t-T}) = a - [\log(y_{i,t-T})]\cdot[1-\exp(-\beta T)]/T$ + regional dummies, where T is 23 years and y_{it} is the ratio of the sector's contribution to state i's gross state product to employment in the sector for that state. The sector for agriculture is omitted because of unreliable data on employment. The regional dummies are for south, midwest, and west. The 8-sector regression with a single value for β comes from iterative, weighted non-linear least squares. Standard errors are in parentheses, $\hat{\sigma}$ is the standard error of estimate.

Table 5: List of European Regions shown in Figures 13 and 14

Germany

1. Schleswig-Holstein
2. Hamburg
3. Niedersachsen
4. Bremen
5. Nordrhein-Westfalen
6. Hessen
7. Rheinland-Pfalz
8. Saarland
9. Baden-Württemberg
10. Bayern
11. Berlin (West)

United Kingdom

12. North
13. Yorkshire-Humberside
14. East Midlands
15. East Anglia
16. South-East
17. South-West
18. North-West
19. West Midlands
20. Wales
21. Scotland
22. Northern Ireland

Italy

23. Piemonte
24. Valle d'Aosta
25. Liguria
26. Lombardia
27. Trentino-Alto Adige
28. Veneto
29. Friuli-Venezia, Giulia
30. Emilia-Romagna
31. Marche
32. Toscana
33. Umbria
34. Lazio
35. Campania
36. Abruzzi
37. Molise
38. Puglia
39. Basilicata
40. Calabria
41. Sicilia
42. Sardegna

France

43. Region Parisienne
44. Champagne-Ardenne
45. Picardie
46. Haute Normandie
47. Centre
48. Basse Normandie
49. Bourgogne
50. Nord-Pas-de-Calais
51. Lorraine
52. Alsace
53. Franche-Comte
54. Pays de la Loire
55. Bretagne
56. Poitou-Charentes
57. Aquitanie
58. Midi-Pyrénées
59. Limousin
60. Rhône-Alpes
61. Auvergne
62. Languedoc-Roussillon
63. Provence-Alpes-Cote d'Azur-Corse[a]

Netherlands

65. Noord
66. Oost
67. West
68. Zuid

Belgium

69. Vlaanderen
70. Wallonie
71. Brabant

Denmark

72. Sjaelland-Lolland-Falster-Bornholm
73. Fyn
74. Jylland

[a]GDP data from Eurostat for Corse were combined with those for Provence-Alpes-Cote d'Azur.

Table 6: Regressions for GDP across Regions of Europe

	(1)		(2)		(3)	
	No other var's.		Country dums.		Country dums., AGRY, INY	
Period	$\hat{\beta}$	R^2 $[\hat{\sigma}]$	$\hat{\beta}$	R^2 $[\hat{\sigma}]$	$\hat{\beta}$	R^2 $[\hat{\sigma}]$
1950-1960	.0106 (.0051)	.06 [.0155]	.0105 (.0038)	.78 [.0077]	.0206 (.0078)	.80 [.0076]
1960-1970	.0367 (.0066)	.39 [.0149]	.0279 (.0036)	.92 [.0057]	.0241 (.0062)	.92 [.0058]
1970-1980	.0035 (.0035)	.01 [.0098]	.0184 (.0049)	.43 [.0078]	.0139 (.0082)	.44 [.0078]
1980-1985	.0953 (.0122)	.60 [.0212]	.0116 (.0048)	.95 [.0077]	.0111 (.0060)	.96 [.0070]
4 periods, β restricted	.0183 (.0029)	--	.0186 (.0021)	--	.0178 (.0034)	--
Likelihood-ratio statistic for equal β's (.05 χ^2 value with 3 df = 7.8):	70.9		13.3		2.6	
p-value:	(.000)		(.004)		(.457)	

Note: 73 observations. The regressions use non-linear least squares and take the form: $(1/T)\cdot\log(y_{it}/y_{i,t-T}) = a - [\log(y_{i,t-T})]\cdot[1-\exp(-\beta T)]/T +$ other variables, where T is the length of the interval and y_{it} is per capita gross domestic product for region i at time t. The other variables in column 2 are country dummies. The breakdown of the sample is 11 regions for Germany, 11 for the U.K., 20 for Italy, 21 for France, 3 for the Netherlands, 3 for Belgium, and 3 for Denmark. The regression in column 3 adds the shares of agriculture and industry in employment at the start of the sub-period (based on a 3-way division of employment into agriculture, industry, and services) for the sub-periods 1950-1960, 1960-1970, and 1970-1980. The regression for the sub-period 1980-1985 includes the shares of agriculture and industry in GDP at the start of the period. The 4-period regression with a single value for β comes from iterative, weighted non-linear least squares. Standard errors are in parentheses, $\hat{\sigma}$ is the standard error of estimate.

Table 7: The North and South of Italy and Great Britain

	Italy		Great Britain	
	North	South	North	South
$\log(y_{1950})$	0.532	-0.385	-0.002	0.003
Growth: 50-60	-0.0014	0.0002	-0.0002	0.0003
$\log(y_{1960})$	0.518	-0.383	-0.004	0.006
Growth: 60-70	-0.0155	0.0080	-0.0027	0.0041
$\log(y_{1970})$[a]	0.363	-0.303	-0.031	0.047
$\log(y_{1970})$[b]	0.404	-0.344	-0.044	0.066
Growth: 70-80	-0.0075	0.0042	0.0021	-0.0031
$\log(y_{1980})$	0.329	-0.302	-0.023	0.035
Growth: 80-85	-0.0011	0.0023	-0.0016	0.0024
$\log(y_{1985})$	0.324	-0.290	-0.031	0.047

Notes: The four northern regions for Italy are Piemonte (no. 23 in Table 5), Valle d'Aosta (24), Liguria (25), and Lombardia (26). The seven southern regions are Campania (35), Abruzzi (36), Molise (37), Puglia (38), Basilicata (39), Calabria (40), and Sicilia (41). The six northern regions for the United Kingdom are North (12), Yorkshire-Humberside (13), North-West (18), West Midlands (19), Wales (20), and Scotland (21). The four southern regions are East Midlands (14), East Anglia (15), South-East (16), and South-West (17). $\log(y_{19xx})$ is the unweighted mean for the indicated regions of the log of per capita personal income, expressed as a deviation from the unweighted mean for the respective country, Italy or Great Britain (Northern Ireland is excluded here). Growth: xx-xx is the unweighted average for the indicated regions of the annual growth rate of per capita personal income, expressed as a deviation from the unweighted average for the respective country.

[a]Data from Molle (1980).

[b]Data from Eurostat.

Appendix Table 1

Unweighted Means and Standard Deviations of Variables

Data for U.S. States

Growth rate of real per capita personal income, $(1/T)\cdot\log(y_{it}/y_{i,t-T})$:

	Mean	σ
1880-1988[a]	.0181	.0045
1880-1900[a]	.0126	.0083
1900-1920	.0138	.0105
1920-1930	.0030	.0140
1930-1940	.0144	.0090
1940-1950	.0492	.0147
1950-1960	.0143	.0065
1960-1970	.0308	.0063
1970-1980	.0150	.0067
1980-1988	.0204	.0141

Log of real per capita personal income, $\log(y_{it})$, where y_{it} is in 1000s of nominal dollars per person divided by the overall CPI (1982 base = 1.0):

	Mean	σ
1880[a]	0.478	.545
1900	0.719	.465
1920	0.995	.327
1930	1.026	.401
1940	1.170	.356
1950	1.661	.244
1960	1.805	.208
1970	2.112	.168
1980	2.262	.150
1988	2.425	.194

Shares of personal income originating in agriculture, $AGRY_{it}$:

	Mean	σ
1880[b]	.307	.184
1900	.273	.150
1920	.211	.120
1930	.134	.087
1940	.122	.084
1950	.117	.087
1960	.058	.050
1970	.040	.040
1980	.020	.019

Appendix Table 1, page 2

	Mean	σ
Regional dummies:		
East	.229	--
South	.292	--
Midwest	.250	--
West	.229	--
Structure variable, S_{it}, for personal income:		
1930-1940	.0164	.0012
1940-1950	.0393	.0020
1950-1960	.0103	.0082
1960-1970	.0254	.0028
1970-1980	.0044	.0026
1980-1988	.0464	.0058
Rates of net in-migration, m_{it}:		
1900-1987	.0034	.0107
1900-1940	.0051	.0119
1940-1987	.0019	.0113
1900-1920	.0107	.0187
1920-1930	-.0002	.0115
1930-1940	-.0086	.0069
1940-1950	.0004	.0140
1950-1960	.0009	.0167
1960-1970	.0009	.0112
1970-1980	.0055	.0123
1980-1987	.0020	.0086
Population density, n_{it}, in 1000s of persons per square mile of total area:		
1880[a]	.0388	.0521
1900	.0559	.0797
1920	.0771	.1145
1930	.0906	.1342
1940	.0935	.1379
1950	.1062	.1553
1960	.1247	.1817
1970	.1416	.2081
1980	.1504	.2099
$HEAT_i$	5033	2116
$\log(HEAT_i)$	8.407	.539

Appendix Table 1, page 3

	Mean	σ
Growth rate of per capita GSP, where nominal GSP is deflated by national price index for GSP, $(1/T)\cdot\log(y_{it}/y_{i,t-T})$:		
1963-1986	.0227	.0050
1963-1969	.0370	.0087
1969-1975	.0159	.0149
1975-1981	.0207	.0146
1981-1986	.0159	.0260
Log of per capita gross state product, $\log(y_{it})$ (y_{it} is in 1000s of 1982 dollars per person):		
1963	2.138	.181
1969	2.360	.155
1975	2.456	.145
1981	2.580	.181
1986	2.659	.142
Structure variable, S_{it}, for gross state product:		
1963-1969	.0282	.0037
1969-1975	.0053	.0060
1975-1981	.0156	.0080
1981-1986	.0158	.0112
Growth rate of sectoral productivity (contribution to real GSP per worker) from 1963 to 1986:		
Construction	-.0223	.0117
Mining[c]	-.0082	.0178
Manufacturing	.0282	.0076
Transportation	.0230	.0047
Trade	.0105	.0033
FIRE	.0006	.0058
Services	-.0053	.0041
Government	-.0062	.0045
Log of sectoral productivity in 1963, $\log(y_{it})$, where y_{it} is real GSP per worker in 1000s of 1982 dollars:		
Construction	2.233	.172
Mining[c]	2.727	.505
Manufacturing	2.213	.224
Transportation	2.631	.076
Trade	2.065	.087
FIRE	3.494	.190
Services	1.996	.118
Government	1.840	.216

Appendix Table 1, page 4

	Mean	σ
Log of sectoral productivity in 1986, $\log(y_{it})$, where y_{it} is real GSP per worker in 1000s of 1982 dollars:		
Construction	3.697	.282
Mining[c]	4.353	.456
Manufacturing	3.737	.119
Transportation	4.300	.109
Trade	3.318	.096
FIRE	4.750	.153
Services	3.325	.114
Government	3.285	.161
<u>Data for European Regions</u>		
Growth rate of per capita GDP[d]:		
1950-1985	--	.0088
1950-1960	--	.0159
1960-1970	--	.0190
1970-1980	--	.0098
1980-1985	--	.0331
Log of per capita GDP:[d]		
1950	--	.395
1960	--	.387
1970[e]	--	.306
1970[f]	--	.334
1980	--	.337
1985	--	.234
Agriculture share of employment:		
1950	.319	.199
1960	.231	.165
1970	.142	.110
Industry share of employment:		
1950	.373	.125
1960	.412	.106
1970	.430	.080

Appendix Table 1, page 5

	Mean	σ
Agriculture share of GDP[g]:		
1970	.076	.051
1980	.050	.035
1985	.045	.030
Industry share of GDP[g]:		
1970	.430	.079
1980	.403	.066
1985	.362	.065
Country dummies:		
Germany	.151	--
Italy	.274	--
U.K.	.151	--
France	.288	--
Netherlands	.055	--
Belgium	.041	--
Denmark	.041	--

[a] 47 observations (excluding Oklahoma).

[b] 46 observations (excluding Oklahoma and Wyoming).

[c] 42 observations (excluding Connecticut, Delaware, Maine, Massachusetts, New Hampshire, and Rhode Island, which have negligible mining).

[d] Levels of per capita GDP for different years are based on non-comparable indexes.

[e] Data from Molle (1980).

[f] Data from Eurostat.

[g] Excluding the three regions for Denmark. For the regressions, the 1980 values for Denmark were approximated from the available data for 1974.

FIGURE 1: GROWTH RATE FROM 1880 TO 1988 VERSUS 1880 PER CAPITA INCOME

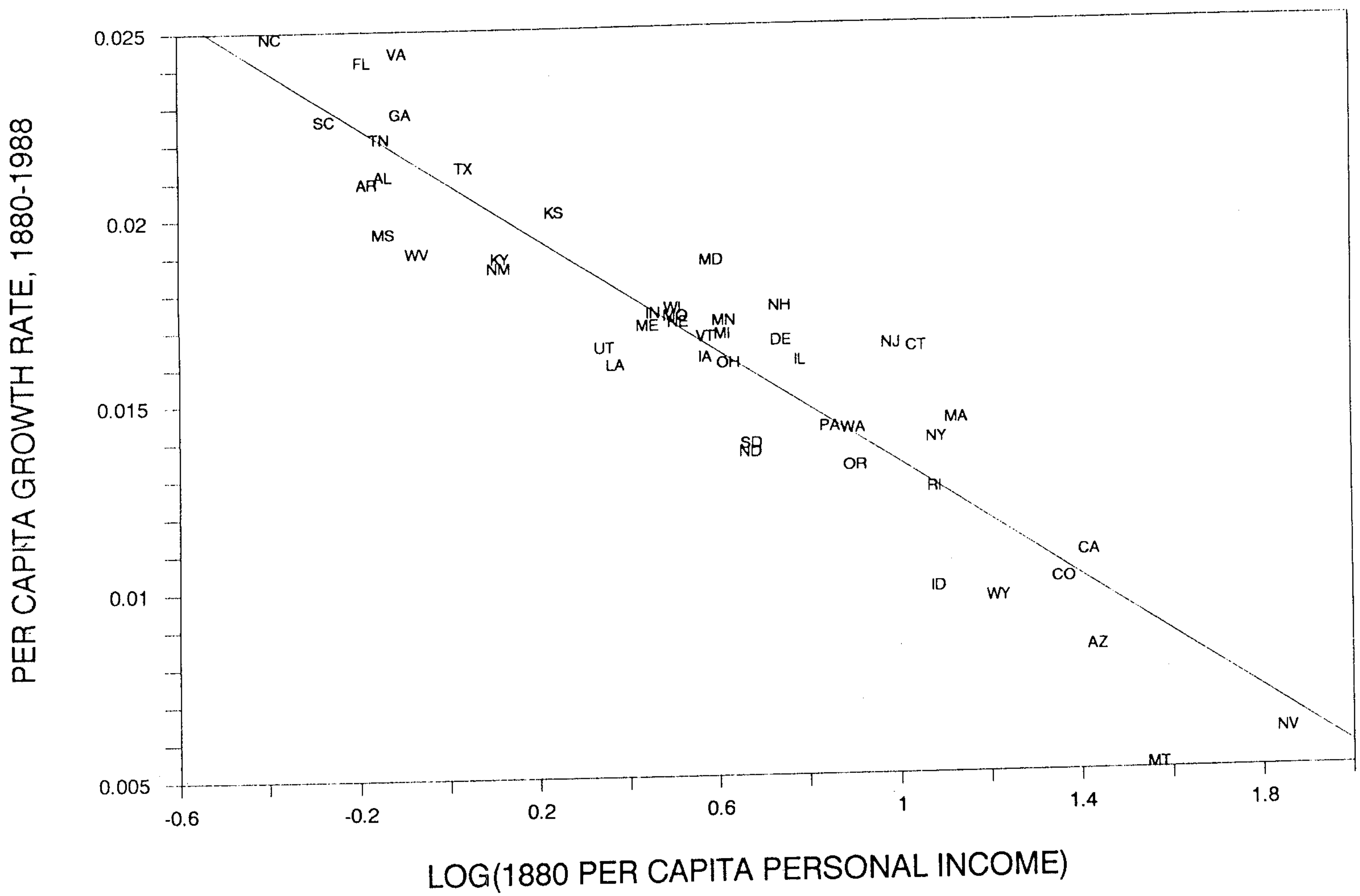

FIGURE 2: GROWTH RATE FROM 1880 TO 1988
VERSUS 1880 PER CAPITA INCOME
FOUR CENSUS REGIONS

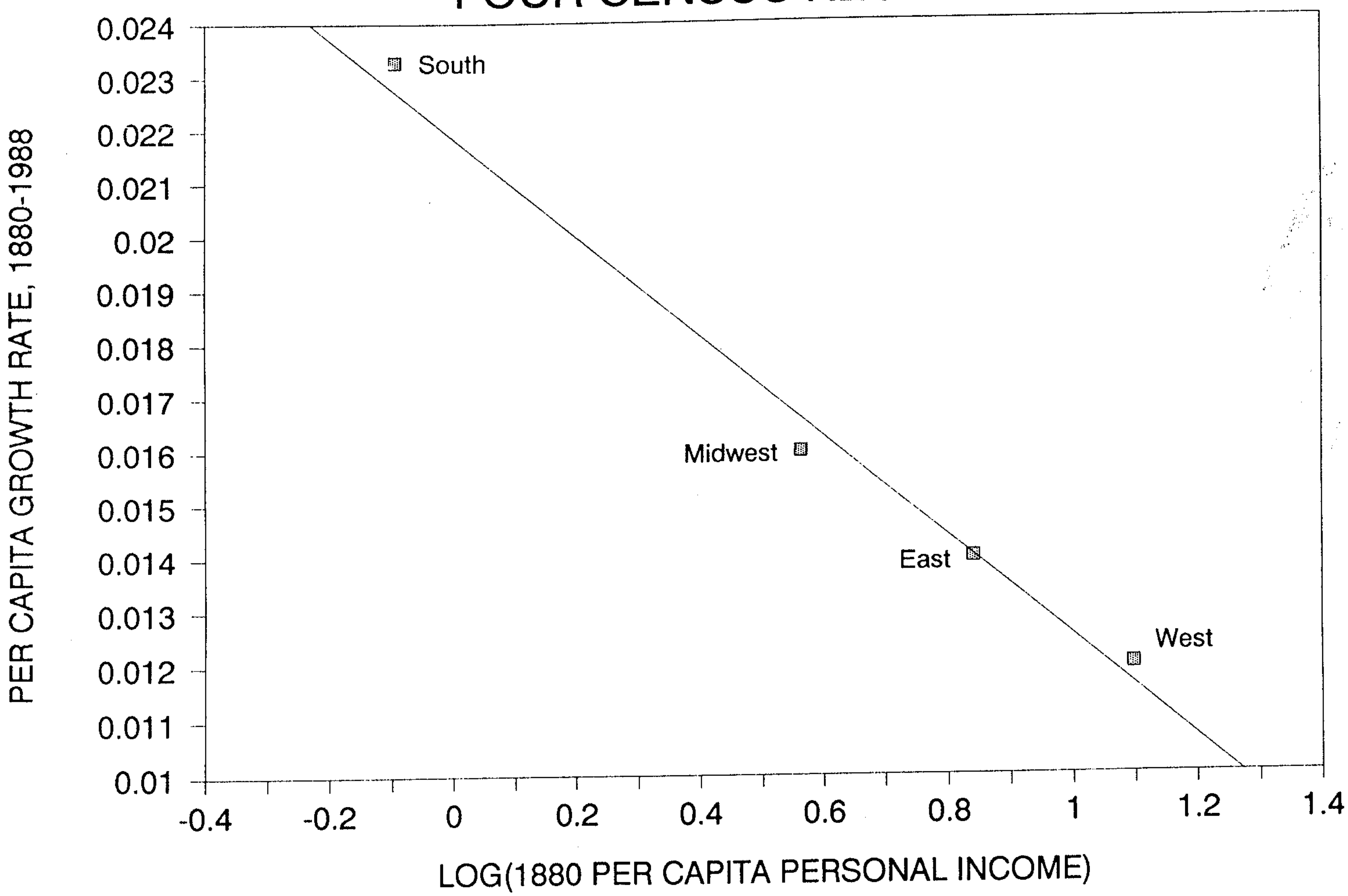

FIGURE 3: PARTIAL RELATION BETWEEN GROWTH RATE
AND LOG OF INITIAL INCOME 1880-1988
HOLDING CONSTANT REGIONAL DUMMIES

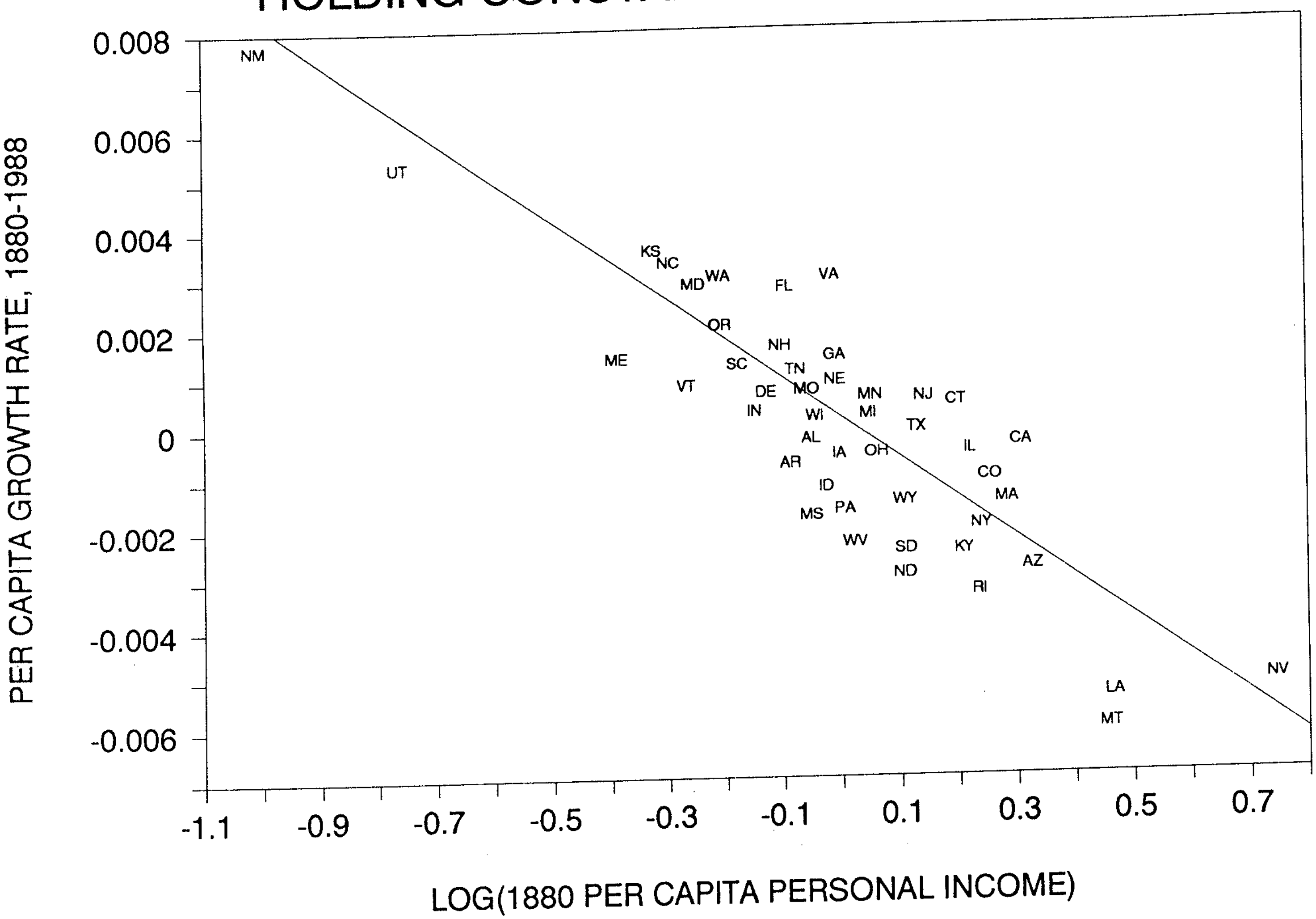

FIGURE 4: DISPERSION, σ_t, OF INCOME PER CAPITA ACROSS U.S. STATES

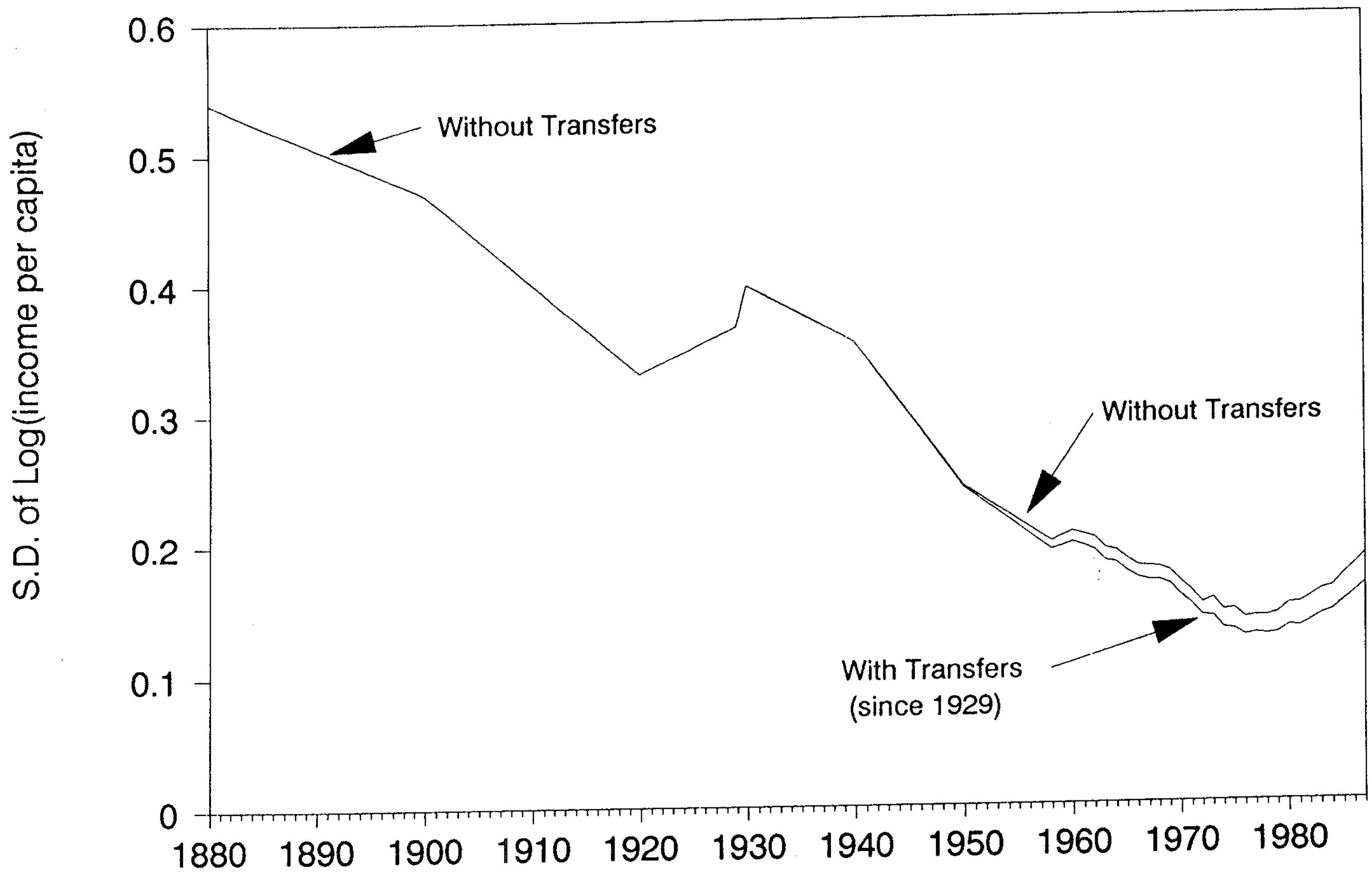

FIGURE 5: DISPERSION, σ_t, OF INCOME PER CAPITA ACROSS FOUR U.S. REGIONS

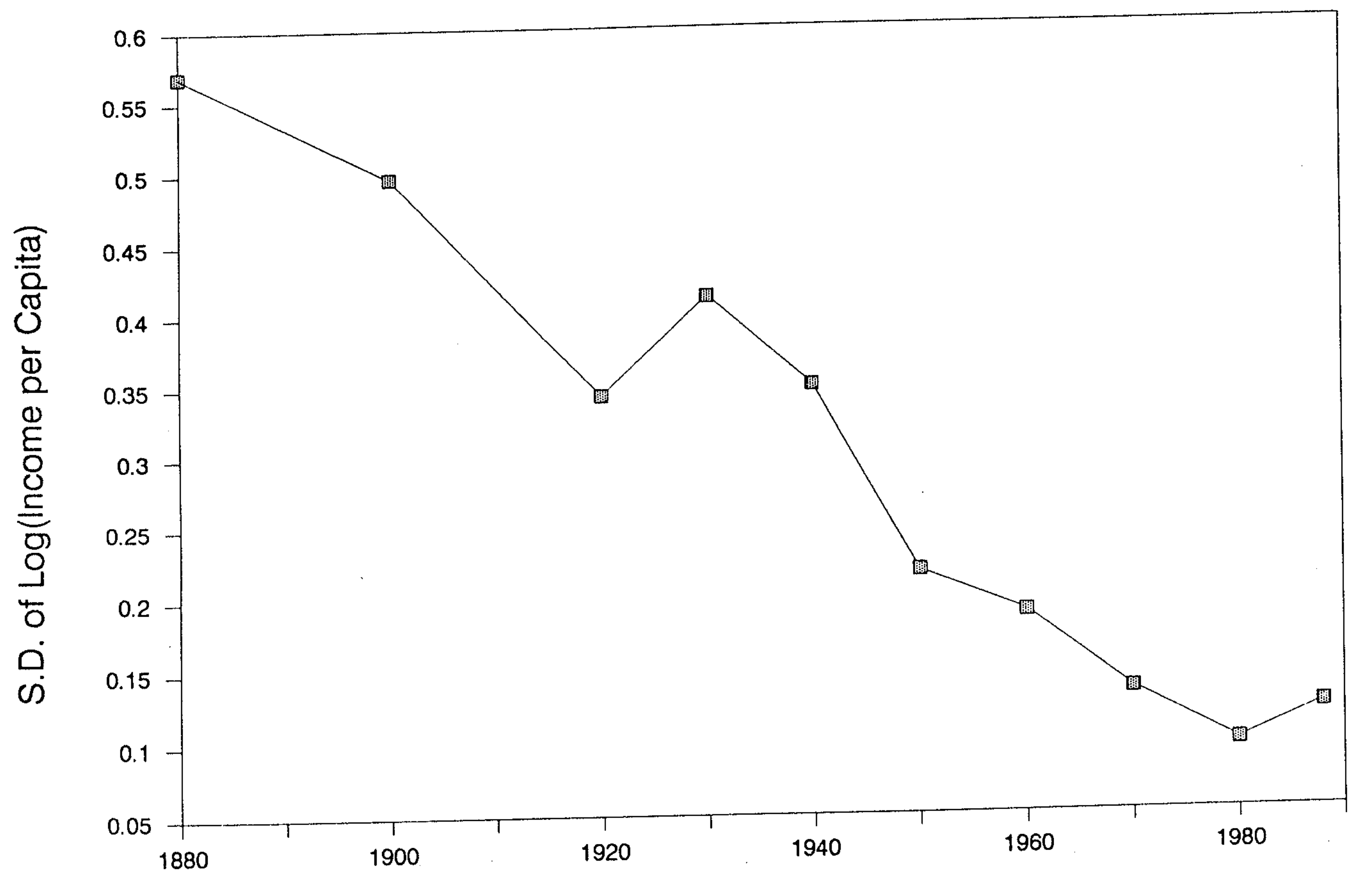

FIGURE 6: PERSONAL INCOME PER CAPITA OVER TIME
FOUR U.S. REGIONS

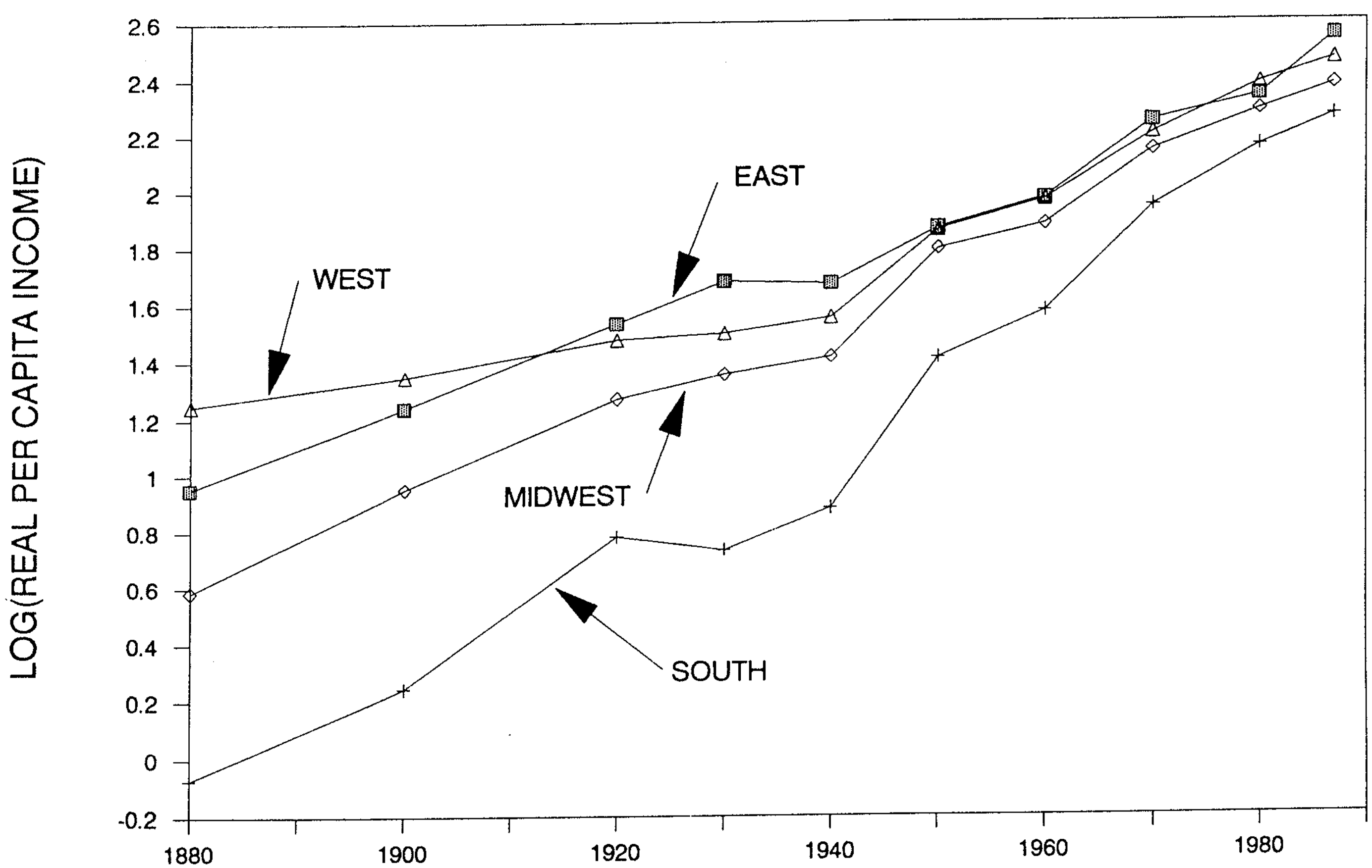

FIGURE 7: DISPERSION, σ_t, OF INCOME PER CAPITA WITHIN FOUR U.S. REGIONS

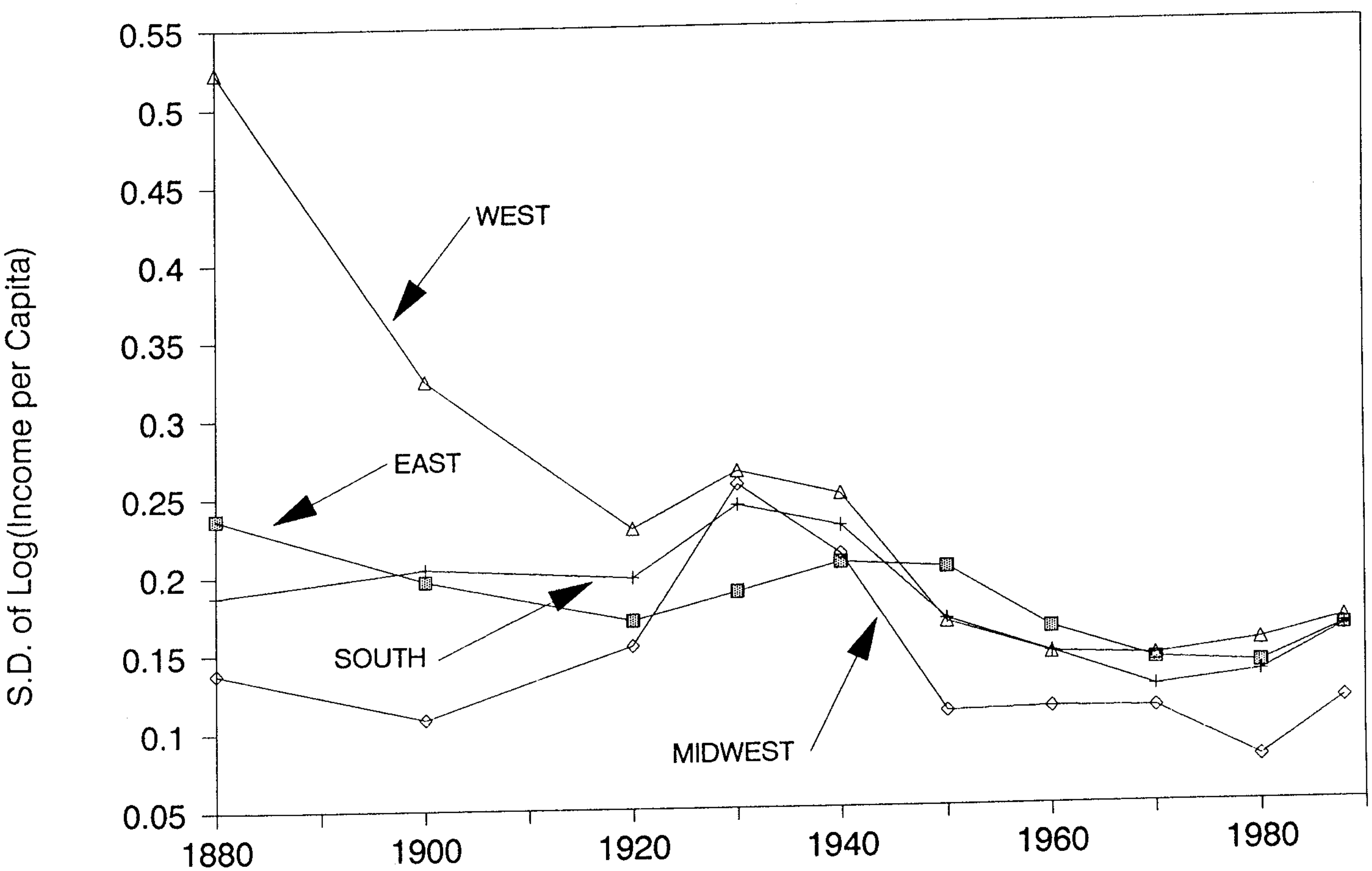

FIGURE 8: MIGRATION VERSUS INITIAL INCOME

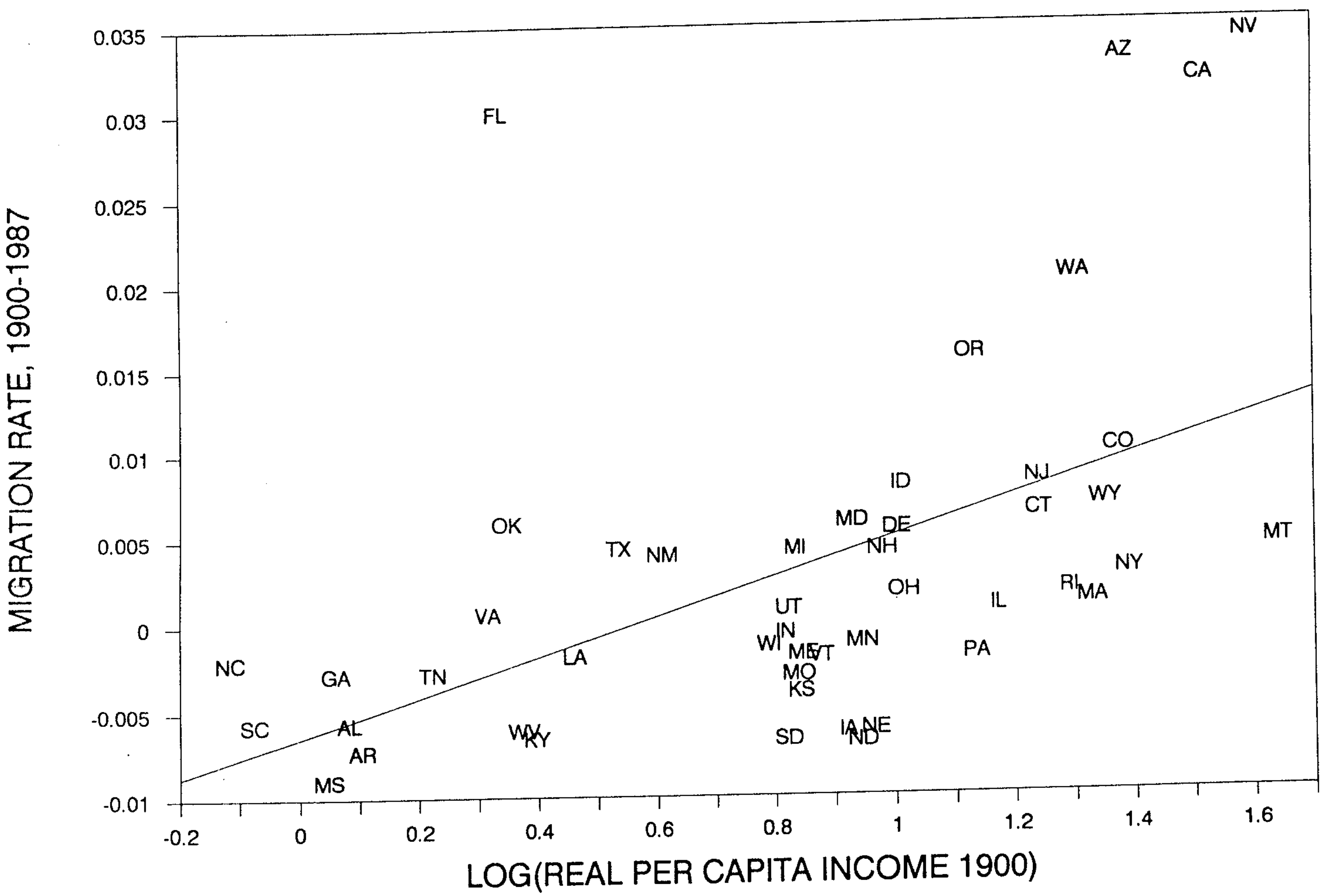

FIGURE 9: PARTIAL RELATION BETWEEN MIGRATION AND INITIAL INCOME

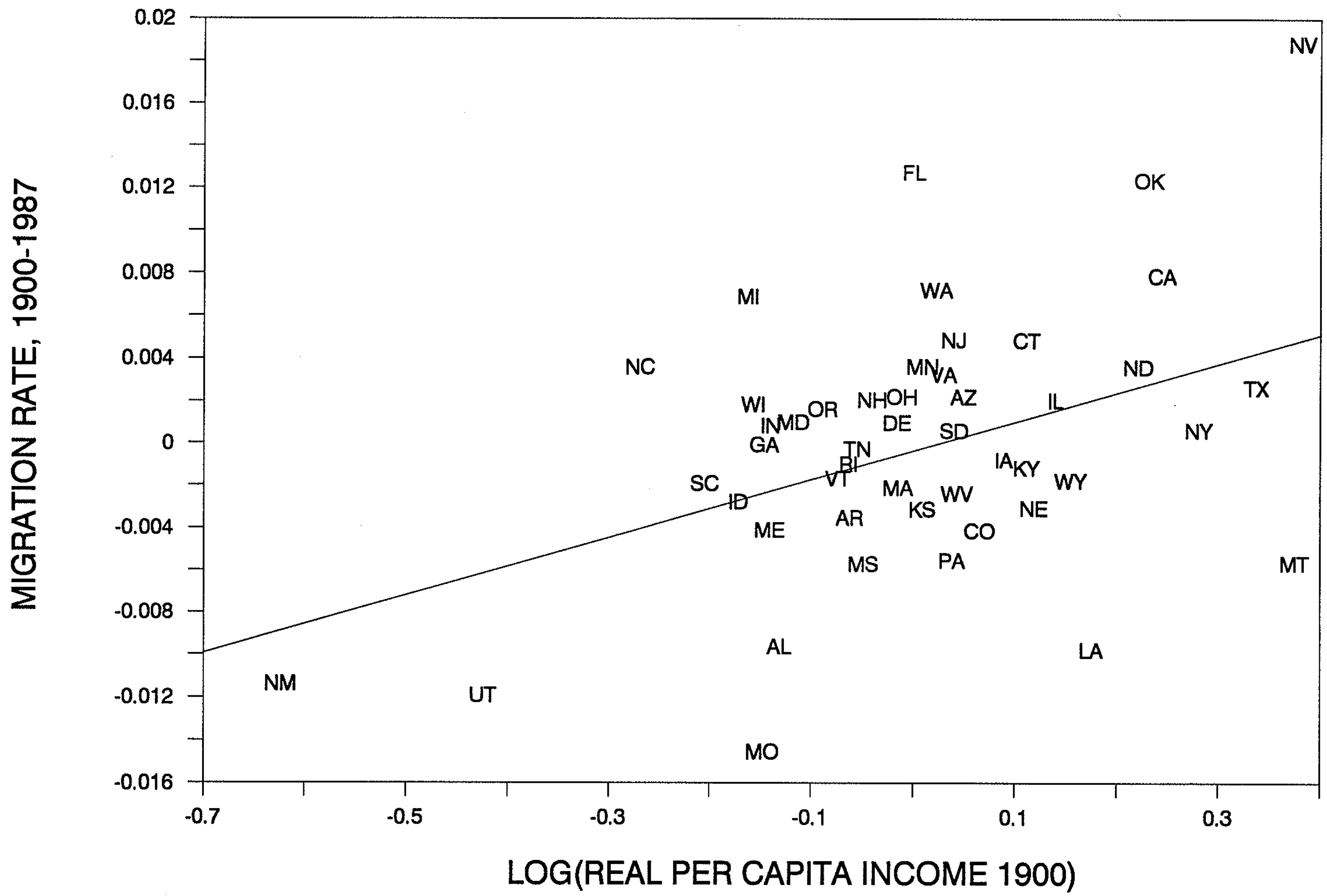

FIGURE 10: PERSISTENCE OF MIGRATION RATES

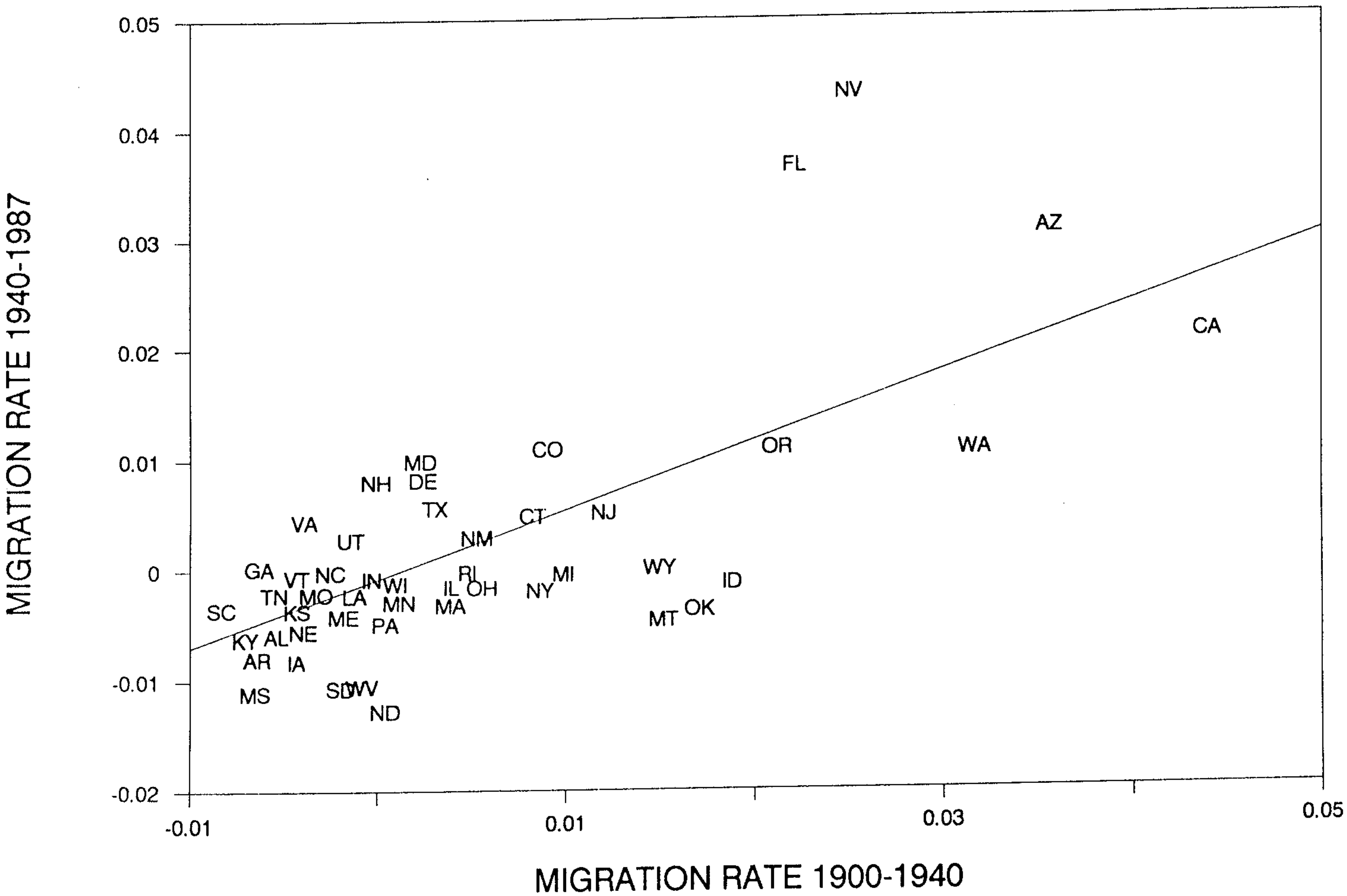

FIGURE 11: GROWTH RATE FROM 1963 TO 1986 VERSUS 1963 PER CAPITA GSP

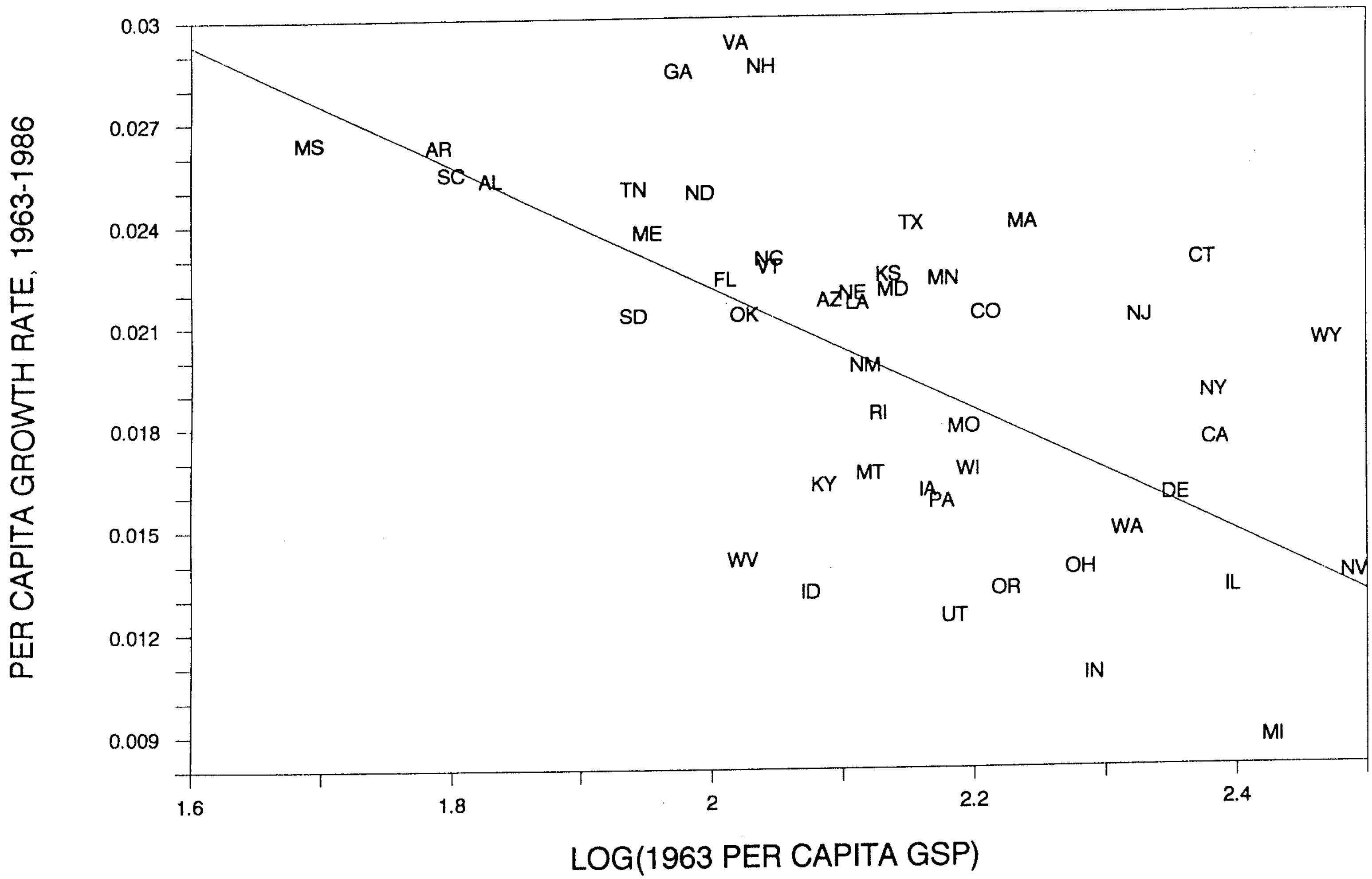

FIGURE 12: DISPERSION, σ_t , OF GSP PER CAPITA ACROSS THE U.S. STATES

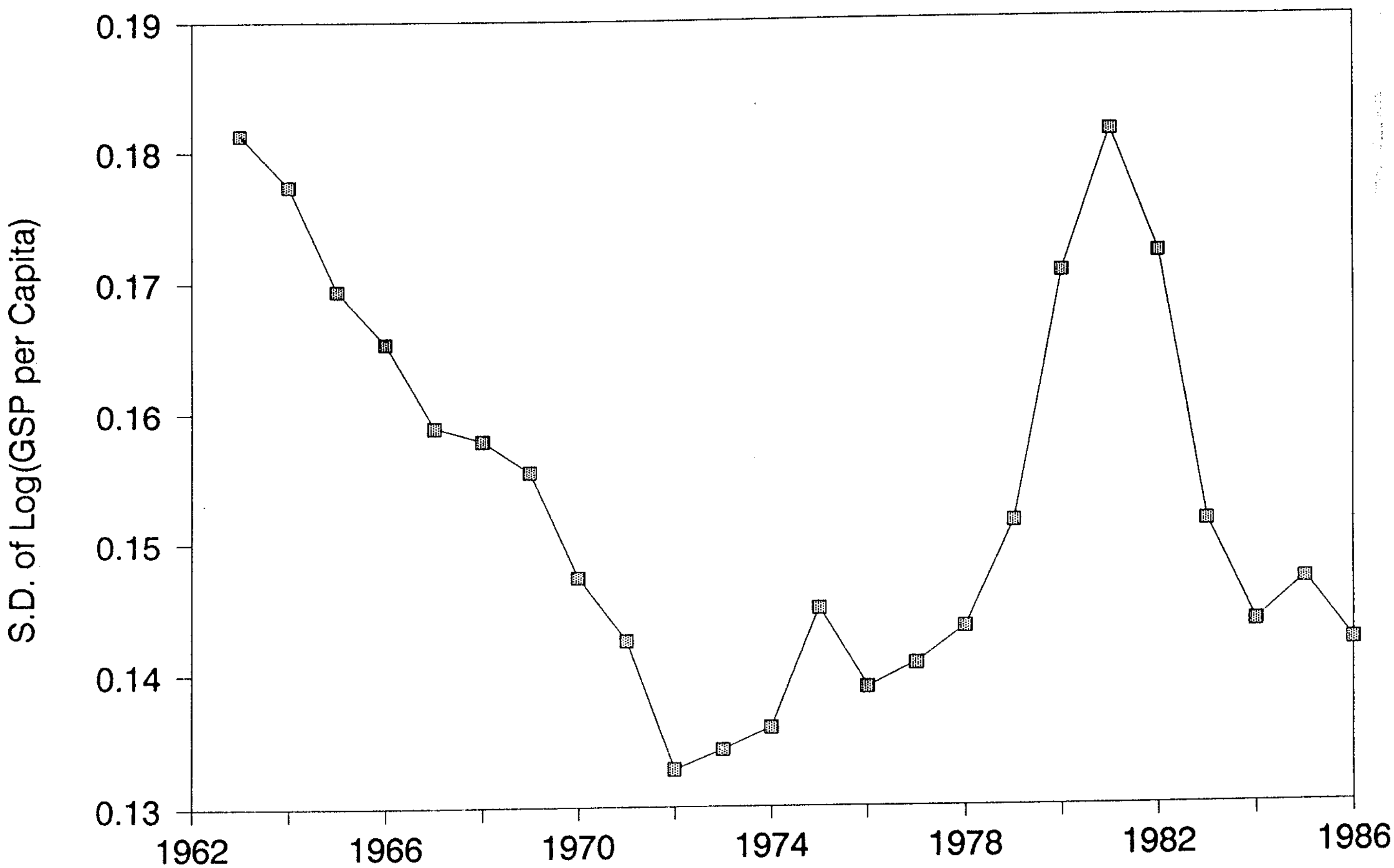

FIGURE 13: GROWTH RATE FROM 1950 TO 1985 VERSUS 1950 PER CAPITA GDP FOR 73 REGIONS IN EUROPE

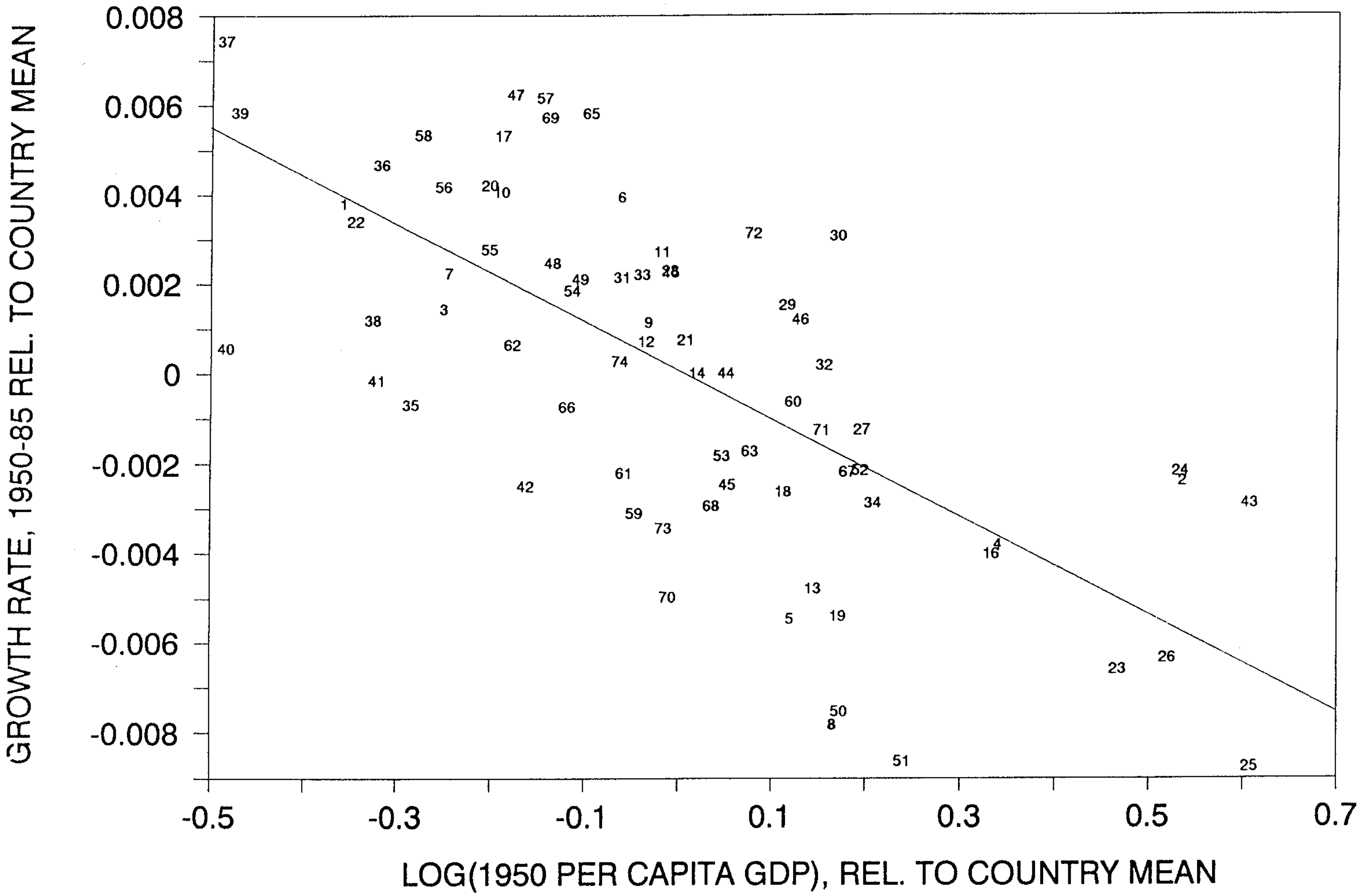

Figure 14

Map of Regions of Europe

(Adapted from Molle [1980, p.20]. See Table 5 for names of regions.)

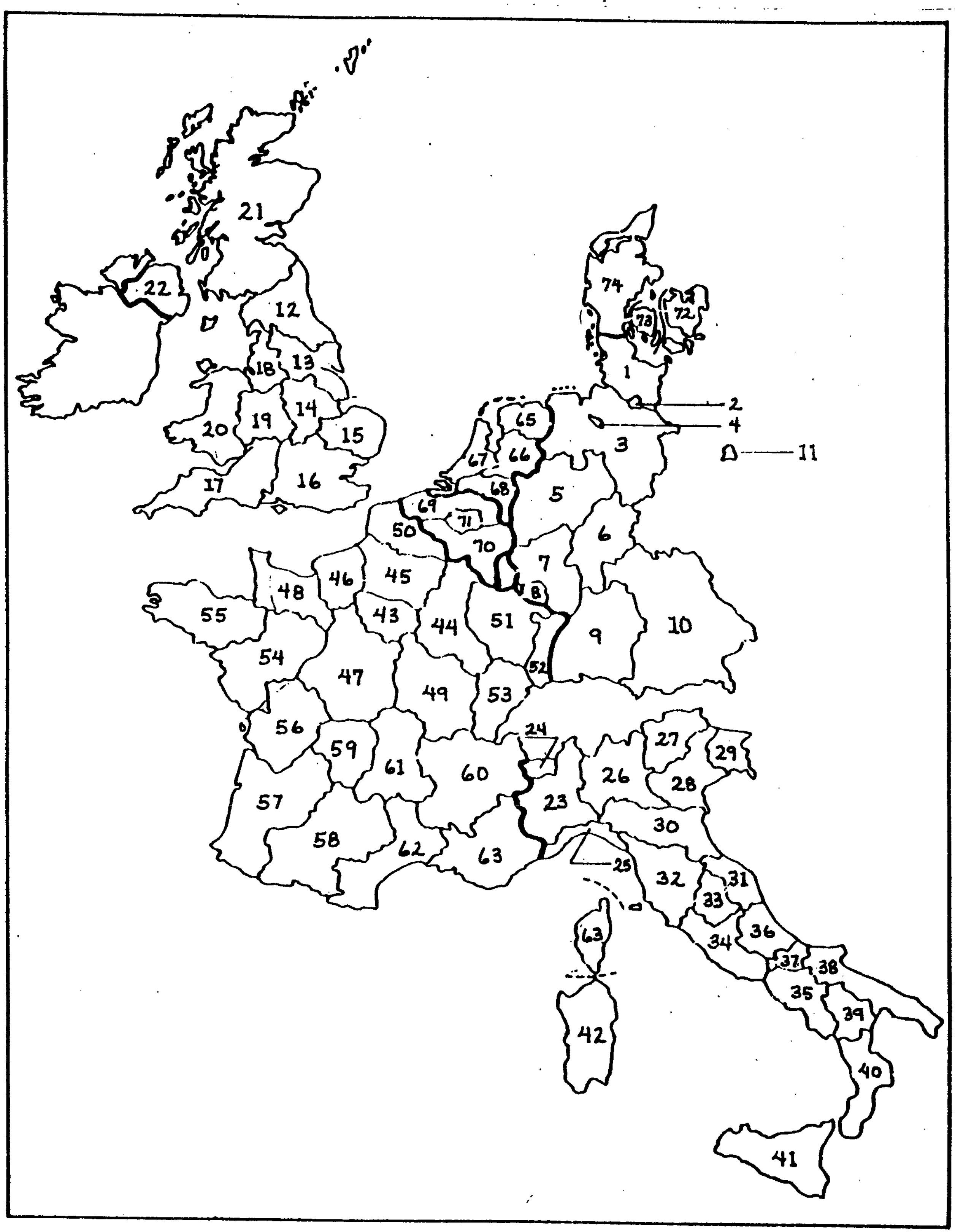

FIGURE 15: DISPERSION, σ_t , OF GDP PER CAPITA ACROSS 73 EUROPEAN REGIONS

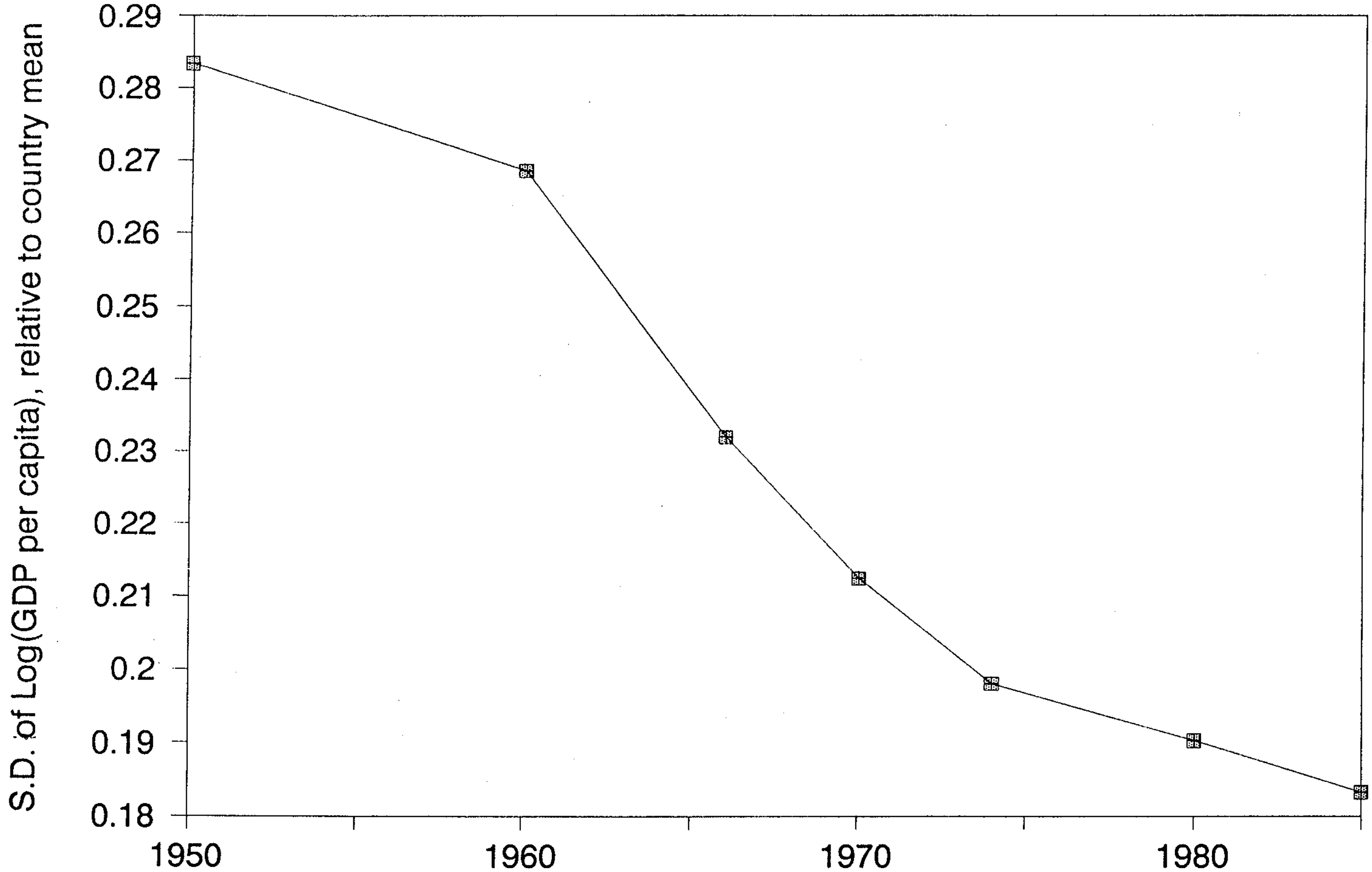

FIGURE 16: DISPERSION, σ_t, OF GDP PER CAPITA WITHIN FOUR EUROPEAN COUNTRIES