INCOME DYNAMICS IN

REGIONS AND COUNTRIES

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1. Introduction

My purpose in this paper is to investigate the dynamics of per capita income in the regions of a given country and, to a lesser extent, in the countries of the world. A key question that I seek to answer is whether the per capita incomes of the regions or countries tend to converge toward parallel balanced growth paths. If an affirmative answer appears to be reasonable, I shall then seek to estimate how rapidly this convergence takes place, how far apart the balanced growth paths are, and what variables determine how high the balanced growth paths are.

The plan of this paper is as follows. Section 2 discusses the implications of growth theory and the answers that the recent empirical growth literature has provided to these questions. It also explains why these answers cannot be given much credence. Section 3 argues that the income dynamics of the regions of a country may be qualitatively different from those for the countries of the world. In particular, per capita incomes could converge toward parallel balanced growth paths for the former while diverging for the latter. Furthermore, to the extent that labor mobility is quantitatively important across a collection of regions within some country, convergence is likely to be much more rapid for them than for countries of the world among which labor mobility is negligible. Section 5 presents empirical results for the 48 contiguous US states. It is hoped that these results can indicate what one might expect to observe for countries at some future date when factors are as mobile across countries as they are now across the states. Section 6 compares these empirical findings with some obtained for a sample of countries. Finally, section 7 offers a few conclusions.

2. Review of the Literature

Robert Solow (1956), David Cass (1965), and Tjalling Koopmans (1965) *inter alios* formulated neoclassical growth theories in which production functions are characterized by diminishing returns to labor and reproducible factors separately and constant returns to them jointly. As a result, their models predict that per capita income can grow in the long run only if exogenous technological improvement keeps the efficiency of labor growing. On the assumption that the efficiency of labor grows at a constant rate, per capita income converges to a unique balanced growth path in their models. In other words, if per capita income is below (above) its level on the balanced growth path, it grows more (less) rapidly than along the balanced growth path. Their models also predict that the height of the balanced growth path toward which per capita income converges is higher, the thriftier and less fertile are the inhabitants of the economy and the more productively they use their resources.

Applying this model to national economies is straightforward on the assumption that factors do not move across national frontiers and countries are identical except for their initial conditions. Given these strong assumptions, neoclassical growth theories predict that countries with high per capita incomes grow more slowly than countries with low per capita incomes. These predictions, however, are grossly inconsistent with the historical experience since 1820. According to Angus Maddison (1995, p. 22), the world's richest country in 1820 had three times the per capita income of the poorest country for which data are available. That ratio rose to 7 in 1870, 11 in 1913, 35 in 1950, 40 in 1973, and 72 in 1992. Table 1 reports his estimates of per capita income for seven country groups over the period 1820-1992, and Table 2 reports growth rates calculated from them. Before 1950, richer groups tended also to grow more rapidly. Only since 1950 has any tendency for rich countries to grow more slowly than poor countries been evident. Even that tendency is confined mostly to Western and Southern Europe and Asia; Latin America and especially Africa have continued to lose ground in relative terms. For large samples of countries over the period since 1960, the correlation between growth rates and initial levels of per capita income is positive; see Robert Barro (1997).

Maddison (p. 20) estimates that world per capita income grew 1.21 percent a year between 1820 and 1992. Growth that rapid is unprecedented in world history and prehistory; see David Landes (1998). Indeed, growth in the previous 10,000 years could not have been even as high as 0.04 percent a year, Maddison's estimate for 1500-1820. The reason is that world per capita income would have had to start at only $10 a year in order to have grown to Maddison's estimate of $565 a year in 1500. Neoclassical growth theory has nothing useful to say about this enormous increase in the growth rate. Nor does it appear to explain why growth rates since 1820 have differed across groups of countries for long periods.

To overcome these shortcomings of neoclassical growth theory, Paul Romer (1986, 1990), Robert Lucas (1988), and Philippe Aghion and Peter Howitt (1992) *inter alios* formulated growth theories that endogenize the steady-state growth rate. Their models essentially replace the assumption of constant returns to all factors and diminishing returns to reproducible factors with the assumption of increasing returns to all factors and constant returns to reproducible factors. These theories, however, also proved to have counterfactual implications. In particular, they imply that growth rates should be highly persistent, rising or falling with investment rates, shares of employment in R&D, and the size and schooling of the population *inter alia*. In fact, growth rates of individual countries evidence little persistence notwithstanding the great persistence in their presumed determinants; see William Easterly, Michael Kremer, Lant Pritchett and Lawrence Summers (1993), Charles Jones (1995), and my 1997a paper. As a result, a consensus has been emerging that theories of endogenous growth are useful primarily for understanding why the world as a whole and countries at the technological frontier can grow in the long run;[1] see Barro (1997, pp. ix-xii, 1-8). Properly used, neoclassical growth theory can then explain cross-country differences in per capita income. Using it properly, however, entails recognizing that technology only gradually diffuses across countries and that countries do differ in more that merely initial conditions. See Landes for a rich account of the many ways in which countries differ.

A large literature has used cross-country data in order to investigate the determinants of how high a given country's balanced growth path is and how rapidly it approaches this path. (Barro, 1997, provides a useful review of this literature.) The basic econometric framework in this literature is the partial-adjustment model

$$(1) \qquad Dy_c = \gamma + \eta(y_c^* - y_c^0), \quad 0 < \eta < 1,$$

where Dy_c is the average growth rate of per capita income over some extended period of time for country c, y_c^* is the steady-state value of the logarithm of per capita income for country c at the beginning of the period, y_c^0 is its actual value, γ is the mean growth rate of the steady-state value of per capita income for the world as a whole and for each

[1] Narayana R. Kocherlakota and Kei-Mu Yi (1997) report evidence of endogenous growth for the United States and the United Kingdom. One interpretation of their finding is that during at least part of each country's sample period, it was the technological leader.

country, and η is the rate at which adjustment occurs. On the assumption that y_c^* is linearly related to a vector x_c of variables and an error term, equation (1) then implies that

(2) $$Dy_c = a + b'x_c - hy_c^0 + x_c,$$

where α is a parameter, β is a vector of parameters, and ξ_c is an error term that is supposed to have a zero mean, a constant variance, and to be uncorrelated across countries. The parameter η measures the rate at which each country converges toward its balanced growth path, accounting not only for the accumulation of reproducible factors but also for the diffusion of technology across countries. Each entry of the parameter vector β equals the product of η and the effect of the corresponding entry of x_c on the height of country *c*'s balanced growth path. Therefore, from knowledge of β and η, one can calculate the effect of x_c on the height of the balanced growth path and hence its unconditional effect on the logarithm of per capita income. Furthermore, one can calculate its predicted effect on the average growth rate Dy_c, conditional on a fixed initial value for per capita income. Unconditionally, however, x_c does not affect country *c*'s growth rate, whose steady-state value γ is exogenous.

This literature reports highly significantly positive estimates of η, which imply convergence toward the balanced growth paths for per capita income of two or three percent a year.[2] Furthermore, these balanced growth paths are estimated to be significantly higher, the more favorable the country's investment climate is, the more educated its population is, and the lower its fertility rate is. As a result, if equation (2) were correctly specified, neoclassical growth theory would receive strong support as a theory of cross-country differences in per capita income.

Unfortunately, equation (2) is likely to be misspecified in empirically important ways. Simultaneity is probably its worst problem. If $\eta > 0$, equation (2) explains the unconditional level of y_c along its balanced growth path as well as the conditional growth rate over the period during which Dy_c is measured. As a result, the same simultaneity problem arises from fitting equation (2) as would arise from regressing y_c^0 directly on x_c; namely, both are biased if per capita income affects any of the variables in x_c. It would be

[2] Barro reports that the convergence rate can be a high as six percent a year for countries with highly educated populations.

astonishing, however, if countries with exogenously higher balanced growth paths did not also choose to educate their populations more, to have fewer children per (more educated, higher-wage) female, and perhaps to have a more favorable climate for investment as well. Furthermore, lagging the variables in x_c is unlikely to help since doing so need not attenuate cross-sectional correlation. For example, educational attainments in 1960, 1975, and 1990 are all likely to be well correlated with the exogenous component of the height of the balanced growth path for per capita income between 1960 and 1990. Finally, finding instrumental variables that are well correlated with the variables in x_c but are not themselves affected by per capita income is a daunting task. Simply lagging the determinants, as is often done in the literature, serves no useful purpose.

A second problem arises because countries with exogenously higher balanced growth paths tend also to have higher initial per capita incomes. As a result, y_c^0 is positively correlated with the error term ξ_c even if x_c is not, biasing the estimates of η and β toward zero. For example, my 1997b paper shows that even if x_c can account for 90 percent of the variance in the heights of the balanced growth paths for y_c, the probability limit of the estimator of η is about $\frac{1}{2}\eta$.

A third problem results from the enormous heterogeneity of countries. This heterogeneity makes the assumption that the parameters of equation (2) are identical across countries simply incredible. In estimating static cross-sectional regressions like those estimated in section 4, it might be reasonable to make this assumption. The reason is that ordinary least squares consistently estimates a weighted average of these parameters under the standard orthogonality restriction on x_t. By contrast, in estimating dynamics, incorrectly assuming identical parameters can result in serious biases and should probably be avoided at all cost; see Kyung So Im, Hashem Pesaran, and Yongcheol Shin (1995).

An alternative to the cross-sectional approach considered above emphasizes the time-series implications of the stock-adjustment model (1). Suppose that only cross-sectional data are available on x_c, the vector of variables useful in explaining the height of country *c*'s balanced growth path for per capita income. In that case, model (1) implies that

(3) $$\Delta y_{ct} = d_c - h_c(y_{c,t-1} - t_{t-1}) + z_{ct};$$

where y_{ct} is the logarithm of per capita income for country c in period t; d_c is a parameter that incorporates the country-specific component of y^*_{ct}, including that explained by x_c; τ_t is the time-specific component of y^*_{ct} (i.e., the common trend of the y^* s); and ζ_{ct} is a stationary error term with a zero mean and finite variance. I give c subscripts to the parameters d_c and η_c in order to allow for cross-country heterogeneity. This unobservable common trend τ_t can be eliminated from equation (3) by averaging equation (3) across the countries and subtracting each member of the resulting equation from the corresponding member of equation (3). The result is

(4) $$\Delta(y_{ct} - \bar{y}_t) = (d_c - \bar{d}) - h_c(y_{c,t-1} - \bar{y}_{t-1}) + (z_{ct} - \bar{z}_t),$$

where $\bar{y}_t$, $\bar{d}$, and $\bar{z}_t$ are the cross-country means of y_{ct}, d_c, and z_{ct}. Because $z_{ct} - \bar{z}_t$ may be serially correlated, one would implement equation (4) empirically in the form

(5) $$\Delta(y_{ct} - \bar{y}_t) = d_c + r_c(y_{c,t-1} - \bar{y}_{t-1}) + \sum_{i=1}^{p} f_{ci}\Delta(y_{c,t-i} - \bar{y}_{t-i}) + u_{ct},$$

where ρ_c is a parameter that is zero or negative depending on whether η_c is zero or positive, δ_c is a parameter with the same sign as $d_c - \bar{d}$, $f_{c1}, f_{c2}, \cdots, f_{cp}$ are parameters arising from the serial correlation of $z_{ct} - \bar{z}_t$, and u_{ct} is a serially uncorrelated error term with a zero mean and finite variance. Estimating equation (5), which takes the form of an augmented Dickey-Fuller regression, then enables one to infer whether the per capita incomes of the countries in the sample converge toward balanced growth paths. Given convergence, one can also estimate how high each country's balanced growth path is and how rapidly convergence toward it occurs.

Andrew Bernard and Steven Durlauf (1995) have applied a variant of this approach to data for individual countries, finding little evidence of convergence. Unfortunately, time-series tests based on one or a few countries for sample periods of even 100 years have little power to reject the null hypothesis of no convergence. Substantial power can be obtained if data for a large number C of countries are pooled and the restrictions $r_1 = r_2 = \cdots = r_C \equiv r$, $f_{11} = f_{21} = \cdots = f_{C1}$, $f_{12} = f_{22} = \cdots = f_{C2}$, $\cdots$, $\phi_{1p} = \phi_{2p} = \cdots = \phi_{Cp}$ are imposed. Doing so, Nazrul Islam (1995) rejects the null hypothesis that ρ = 0 at very small levels and estimates that the convergence rate is nearly 10 percent a year. Bernard and Jones (1996a, 1996c) also reject the null hypothesis for the labor and total factor productivities in a panel of 14 OECD countries.

Unfortunately, these restrictions can be easily rejected and imposing them strongly affects the results; see Kevin Lee, Pesaran, and Ron Smith (1998). Following Andrew Levin and Chien-Fu Lin (1993), Georgios Karras and I (1996) used a procedure that imposes only the restriction $r_1 = r_2 = \cdots = r_C$, also obtaining a strong rejection of ρ = 0. Although this restriction is valid under the null hypothesis that all the ρs are zero, the Monte-Carlo simulations of G.S. Maddala and Shaowen Wu (1997) suggest that appreciable size distortions can result from using this procedure. Finally, application of the procedure of Im, Pesaran, and Shin, which does not impose any of these restrictions, yields little evidence that the ρs differ from zero; see Lee, Pesaran, and Smith.

The cross-sectional approach has been widely used for testing whether the per capita incomes of regions converge toward their balanced growth paths and, if so, to estimate how rapidly the convergence takes place.[3] Some examples are Barro and Xavier Sala-i-Martin (1991; 1992a, 1992b; 1995, Ch. 11) for regions of the United States, Europe, and Japan; Damien Neven and Claudine Gouyette (1995) and Abraham Filip and Paul Van Rompuy (1995) for regions of Europe; Elias Carayannis and Rajiv Mallick (1996) for regions of Canada; and Helmut Hofer and Andreas Worgotter (1997) for regions of Austria. This literature typically finds slow convergence not materially faster than that found for countries. This finding is suspect for two reasons, however. First, the estimates may to be severely biased for the reasons discussed above. Second, because the theories that motivate the empirical work assume no labor mobility, the estimates may be difficult to interpret if labor is in fact highly mobile across the regions. Although the assumption of no labor mobility may be reasonable for countries, it is clearly inappropriate for the regions within many countries. The next section works out the implications of interregional labor mobility in some simple theoretical models.

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[3] These questions have also been addressed using panel methods; e.g., Bernard and Jones (1996b) and Evans and Karras (1996b). They report strong evidence in favor of convergence. Unfortunately, these findings may be unreliable because parameter homogeneity is imposed.

3. Theoretical Implications of Labor Mobility

Typically, the growth literature assumes closed economies. Although that assumption might be defensible for countries, it is implausible for the regions within a country, especially one like the US in which many factors are highly mobile and the barriers to trade in goods and transportation costs are low.

The purpose of this section is to analyze growth within the regions of a country characterized by a high degree of factor mobility. To that end, I assume that labor and perhaps some other factors are costlessly mobile within a country consisting of S regions. Some or all of the goods produced in these regions may also be costlessly tradable interregionally. Of course, not all factors and goods can be costlessly mobile since regions would then cease to be well defined economically. I further assume that there are more immobile factors than mobile factors and goods in each region. This assumption prevents trade in factors and goods from equalizing the prices of the immobile factors across the regions.

I assume that the value Q_{st} of the output produced in region s during period t is related to the labor input N_{st} and the vector X_{st} of the other mobile inputs as follows:[4]

(6) $$Q_{st} = F(N_{st}, X_{st}; \theta_{st}) ,$$

where θ_{st} is a vector of region- and time-specific parameters. Differences in the parameters in (6) reflect differences in inputs of immobile factors across regions and over time as well as differences in product mix and the effects of changes in relative product prices. The function $F(\bullet)$ is assumed not only to be increasing in the mobile factor inputs but also to be strictly concave. If it were convex instead, locating all of the mobile factors in a single region would generally be optimal since marginal products would then be nondecreasing in scale and one region would have higher productivity than the rest.[5] Because complete regional specialization is not observed, however, strict concavity must therefore set in at some scale well short of the entire country's supply of mobile factors.[6]

[4] Of course, the vector X_{st} may include physical capital.

[5] Of course, costs of transport for either factors or goods can sustain nonspecialization in the presence of a moderate degree of convexity. See Paul Krugman (1991) for an example.

[6] Here I am defining *region* broadly enough for the statement to hold. Agglomeration economies are well known to be important in explaining the existence and size of cities.

For simplicity, I assume strict concavity at every scale. Without loss of generality, I can further assume that in equilibrium, the labor input in every region is positive.

Labor receives the same remuneration in every region by assumption. If labor is also paid its marginal product, equation (6) implies that

(7) $$q_{st} \equiv Q_{st}/N_{st} = w_t/\alpha_{st},$$

where w_t be the wage rate and α_{st} is the elasticity of the value of output with respect the labor input in region s during period t. The elasticity α_{st} should be expected to vary across regions because of differences in product mix and over time because of changes in either relative prices or shifts in technologies. Equation (7) therefore implies that labor productivity is likely to vary across regions even though labor is paid at the same rate everywhere and to vary over time as well.[7]

Let y_{st} be the logarithm of productivity and $\bar{y}_t$ be the cross-regional mean of the y_{st}s. On the assumption that the elasticities $a_{1t}, a_{2t}, \cdots, a_{St}$ are covariance stationary, the above analysis indicates that $y_{1t} - \bar{y}_t, y_{2t} - \bar{y}_t, \cdots, y_{St} - \bar{y}_t$ should also be covariance stationary with nonzero means and positive finite variances.[8] Furthermore, the cross-regional variance of the logarithm of labor productivity, defined by

(8) $$V_t = \tfrac{1}{S}\sum_{s=1}^{S}(y_{st} - \bar{y}_t)^2 ,$$

should fluctuate around a positive mean with a positive finite variance.

Thus far, I have assumed that labor is costlessly and immediately mobile. As a result, labor always receives the same remuneration in every region. Suppose instead that labor moves in response to differences in remuneration but the response is distributed over time. Rather than modeling the movement of labor as the rational investment activity of households subject to moving costs, I merely posit an *ad hoc* adjustment equation. Specifically, the labor input in each region adjusts according to[9]

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[7] These implications are similar to those of Jennifer Roback (1982) in a similar model.

[8] Jaume Ventura (1997) has formulated a growth model in which factor prices are equalized across regions. His model implies that their labor productivities can appear to be converging even though growth in each region is endogenous. This result stems in large part from the fact that elasticities $\alpha_1, \alpha_2, \cdots, \alpha_S$ in the model fall toward zero over time if the growth rates of the regions are bounded above zero. As a result, each region's share of wage income in the value of output also approaches zero. Because the wage share has not shown any tendency to vanish in recorded history, this model is not a promising one for interpreting data.

[9] An alternative formulation posits that $\Delta\ln(N_{st}/N_t)$ is proportional to the gap between the logarithm of the wage rate region s during period t and the mean for all regions in the country. Olivier Blanchard and

(9) $$\Delta \ln N_{st} = \Delta \ln N_t + n[\ln N^*_{st} - (\ln N_{s,t-1} + \Delta \ln N_t)], \quad 0 \le n \le 1,$$

where $N^*_{1t}, N^*_{2t}, \cdots, N^*_{St}$ are the efficiency units of labor that would be employed in regions 1, 2, ···, S, were labor instantaneously mobile and N_t is the aggregate labor input. The story underlying equation (9) is that $N^{\circ}_{st} \equiv N_{s,t-1} \exp(\Delta \ln N_t)$ workers find themselves in region s at the beginning of t when wage rates are realized. (The growth rate of each region's labor input is assumed to be $\Delta \ln N_t$ in the absence of migration.) N^*_{st} workers would choose to supply labor in region s, were moving costless. The fixed cost of moving differs across workers, leading some workers to move and others to stay. The distribution of costs each period is assumed to be such that it always optimal to eliminate a fraction ν of the gap between $\ln N^*_{st}$ and $\ln N^{\circ}_{st}$. Consequently, the larger the parameter ν is, the more rapidly labor migrates in response to regional shocks. So long as ν > 0, however, labor is completely mobile in the long run in the sense that the mean of $\ln N_{st} - \ln N^*_{st}$ is zero.

For simplicity, I assume that labor is the only mobile factor and that the elasticity of output with respect to the labor input does not vary across regions or over time. Therefore, the relationship (6) takes the Cobb-Douglas form

(10) $$Q_{st} = N^a_{st} \exp(b_t + a_{st}), \quad 0 < a < 1,$$

where α is the common value of the elasticity of output with respect to the labor input. The quantity b_t+a_{st} is the total factor productivity of region s during period t. It is decomposed into two terms, the first of which is b_t, the cross-regional mean. By definition, then, $\sum_{s=1}^{S} a_{st} = 0$ every period. The second term a_{st} incorporates at least three effects: the idiosyncratic changes in the supplies of the immobile factors in region s; the idiosyncratic changes in the efficiency of those factors induced by technological change; and the idiosyncratic effects of changes in relative prices, given that regions have different product mixes. I assume that a_{st} is difference stationary at worst.

On the assumption that labor is paid its marginal product,

$$aN^{a-1}_{st} \exp(b_t + a_{st}) = aN^{a-1}_{1t} \exp(b_t + a_{1t})$$

Lawrence Katz (1992) have shown that the implications of this formulation are similar to the one here. I adopt equation (9) because it is more tractable and less *ad hoc* than theirs.

so that

(11) $$\ln(N_{st}^{*} \ / \ N_{1t}^{*}) = (a_{st} - a_{1t}) \ / \ (1 - a)\ .$$

Hence,

(12) $$\ln(N_{u}^{*} \ / \ N_{t}) = \frac{a_{u}}{1-a} - \ln \sum_{j=1}^{s} \exp\left(\frac{a_{m}}{1-a}\right)$$

since $\sum_{j=1}^{S} N_{jt}^{*} = N_{t}$. Substituting equation (12) into equation (9) and rearranging produces

(13) $$\ln(N_{u} \ / \ N_{t}) = (1-n)\ln(N_{u,t-1} \ / \ N_{t-1}) + \frac{na_{u}}{1-a} - n \ln \sum_{j=1}^{s} \exp\left(\frac{a_{m}}{1-a}\right)$$

It then follows from equations (13) and (10) that

(14) $$y_{st} - \overline{y}_{t} = (1-n)(y_{s,t-1} - \overline{y}_{t-1}) + (1-n)\Delta a_{st}\ .$$

In the complete absence of mobility (ν = 0), equation (14) implies that $y_{st} - \overline{y}_{t}$ moves one-for-one with the idiosyncratic shock to total productivity in region *s*, exhibiting exactly as much persistence as the shock itself. With some mobility (0 < ν < 1), the contemporaneous effect of the idiosyncratic shock is shrunken by the factor 1-ν, and the persistence of its effect is less than its own. Indeed, even a permanent change in a_{st} affects $y_{st} - \overline{y}_{t}$ only temporarily. With perfect mobility (ν = 1), the idiosyncratic shock ceases to affect $y_{st} - \overline{y}_{t}$ even temporarily. Finally, so long as labor is perfectly mobile in the long run (ν > 0), $y_{st} - \overline{y}_{t}$ should be mean stationary even if the idiosyncratic shocks to total productivity in region *s* are difference stationary.

Equation (14) has two other implications worth spelling out. First, the cross-regional variance of the logarithm of output per worker has a constant mean around which it fluctuates. As a result, one should not expect it to drift downward systematically unless the parameters, which are here assumed to be constant, change systematically. Second, the cross-regional variance is lower, the more mobile labor is. For example, if every region's idiosyncratic shock is a first-order autoregression with the autoregressive parameter π and innovation variance ψ^2, the mean of the cross-regional variance is

$y^2/(1-p^2)$ when labor is completely immobile and 0 when labor is perfectly mobile.[10] Downward drift in the cross-regional variance therefore suggests increasing mobility of factors over time.

At least since Heckscher and Ohlin, economists have widely appreciated that trade in goods can be a substitute for trade in factors. Paul Samuelson (1953) formalized this intuition in his famous factor-price equalization theorem. It is reasonable, then, to expect a reduction in barriers to trade or in transportation costs to have effects similar to those of an increase in labor mobility. See Dan Ben-David (1993) for evidence supporting this conjecture.

Ideally, one would like to test the above theories using measures of labor productivity. In practice, one must typically use measures of per capita income instead. Using the latter, however, raises two issues: households choose the per capita labor supplied to the market; and they choose where to live. The endogeneity of the per capita labor input creates no problem if its long-run supply is vertical. In that case, it fluctuates around a constant mean along the balanced-growth path since the income and substitution effects exactly cancel each other. By contrast, if its supply is not vertical, it diverges across regions. I follow the literature here in assuming that its long-run supply is vertical even though some fairly convincing evidence exists against the assumption; see Evans and Karras (1997b).

The endogeneity of residence may also create problems. For example, suppose that the supply of inputs can be completely divorced from the choice of where to live. In that case, everyone might live in California but own labor and other factors located elsewhere. Dividing the output produced in a given state by the number of individuals residing there would then produce infinite values for every state except California. This problem can be overcome by putting only the income received by the residents in the numerator. The size of the resulting series, however, then partly reflects the effect of income and wealth on the demand for regional amenities. For example, Florida and Arizona, whose winters are warm, might attract relatively wealthy retirees and high-income workers. In practice, however, individuals often live near where their inputs are

[10] The zero value is an artifact of the assumption that the αs are identical across regions and constant over time as the first example considered above makes clear.

supplied. Labor, by its very nature, requires the presence of its owner. For the most part, the same is true of human capital. Physical capital may also tend to be supplied locally since it is typically more productive when managed by its owner than when managed by an employee. The evidence reported by Gur Huberman in this volume suggests that even marketed claims to capital are overrepresented in the portfolios of households that live near the underlying capital. Finally, if households are life-cycle savers, their assets incomes should be cointegrated with their wage incomes. As a result, if per capita labor income is stationary, per capita asset income is also stationary; see Evans and Karras (1997a) for proof. All in all, then, the points made in this section are likely to apply to per capita income as well as labor productivity.

I conclude this section with a discussion of the implications of the accumulation of immobile reproducible factors. Barro, Mankiw, and Sala-i-Martin (1995) have shown that the analysis of such accumulation is similar to its closed-economy counterpart if the mobile factors experience diminishing returns in every region as is assumed here. In particular, accumulation generates additional dynamics in $y_{st} - \bar{y}_t$. The additional dynamics are more persistent, the more elastic the production functions are with respect to these factors. Clearly, the dynamics are likely to be appreciably less persistent for regions than for countries since the class of immobile reproducible factors is smaller for the former than for the latter. Except for this last result, then, allowing for the accumulation of immobile reproducible factors does not alter any of the conclusions of this section.

4. Empirical Analysis for the US States

Many goods and factors have long been quite mobile across the US states; see Sukkoo Kim (1997). State data on per capita income should therefore prove useful in exploring the ideas developed in section 2.

In carrying out this exploration, I use annual data on per capita factor income (personal income less transfer payments) spanning the period 1929-1996 for each of the contiguous US states.[11] I obtained the underlying data from a CD-Rom provided by the

[11] In an earlier paper, I also considered per capita personal income, per capita wage income, and per capita asset income. The results were similar in all cases to those reported here.

Bureau of Economic Analysis (RCN-0128) and from the BEA's web page. I have not bothered to deflate the data since state-specific deflators are unavailable. (Deflating by a common national price index would merely subtract the same value from both y_{st} and $\bar{y}_t$ in $y_{st} - \bar{y}_t$.)

4.1 Is There Convergence?

I first investigate whether the per capita factor incomes of the contiguous US states converge toward parallel balanced growth paths; i.e., whether $y_{st} - \bar{y}_t$ is mean stationary. For this purpose, I employ the procedure formulated by Im, Pesaran, and Shin. It entails using ordinary least squares to fit equation (5) to the data for each state *s* and then calculating the following test statistic

$$(15) \qquad \sqrt{S}\sum_{s=1}^{S}(\hat{t}_s - m_{DF}) / s_{DF},$$

and comparing it to the critical values from the standard normal distribution. In expression (15), $\hat{t}_s$ is the *t*-ratio for the estimate of ρ_s, and μ_{DF} and σ_{DF} are the mean and standard deviation of the appropriate Dickey-Fuller distribution. A sufficiently negative test statistic leads to rejection of the null hypothesis

$$H_0: (\rho_1 = 0 \wedge \rho_2 = 0 \wedge \cdots \wedge \rho_S = 0) \wedge (\delta_1 = 0 \wedge \delta_2 = 0 \wedge \cdots \wedge \delta_S = 0)$$ [12]

in favor of the alternative hypothesis

$$H_1: \rho_1 < 0 \wedge \rho_2 < 0 \wedge \cdots \wedge \rho_S < 0.$$ [13]

The reason is that if H_0 is true, the *t*-ratios are independent of each other and each follows a Dickey-Fuller distribution.

Table 3 reports summary statistics on the *t*-ratios obtained by fitting equation (5) to the logarithm of per capita factor income. For purposes of comparison, the last column

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[12] I impose the condition that the δs are all zero for two reasons. First, I think that it is highly implausible that states can have endogenously different trend growth rates. Second, the resulting test is more conservative than one that left the δs unrestricted. This restriction still allows per capita incomes to diverge across regions so long as they only wander apart rather than trend apart.

[13] If any ρ is negative, all must be negative, and if any is zero, all must be zero; see my 1998 paper for proof. In other words, H_1 is equivalent to $\rho_1 < 0 \vee \rho_2 < 0 \vee \cdots \vee \rho_S < 0$. Truth of H_0, however, does not rule out the possibility that some proper subsets of the regions have per capita incomes, whose pairwise logarithmic differences are stationary. To formulate a test for the presence of such subsets when no evidence for H_1 can be found is a difficult statistical problem, well beyond the scope of this paper. Fortunately, the issue is moot since H_0 can be readily rejected with the data analyzed here.

of the table reports analogous statistics for the appropriate Dickey-Fuller distribution. Clearly, the empirical distribution for the *t*-ratios is shifted considerably to the left relative to the Dickey-Fuller distribution. This result suggests a pronounced tendency of the logarithm of every series to converge toward a common national trend. Formal confirmation is provided by the test statistic, which is -6.64. The null hypothesis of difference stationarity can therefore be rejected at any reasonable significance level.[14]

4.2 How Rapid Is the Convergence?

Given the strong evidence for the convergence of the series toward a common trend, it makes sense to estimate how rapidly the convergence takes place. Estimation of convergence rates is complicated by two considerations, however. First, unless $y_{st} - \bar{y}_t$ is a first-order autoregression (i.e., $p = 0$), y_{st}'s convergence rate depends on the horizon considered. I therefore focus on the asymptotic convergence rate, which prevails at arbitrarily long horizons. It is $1-\lambda_s$, where λ_s is the dominant root of the polynomial

$$(16) \qquad z^{p+1} - (1 + r_s)z^p + (1 - z)\sum_{i=1}^{p} f_{si} z^{p-i} \; .$$

In many applications, this measure provides an excellent approximation to the convergence rates prevailing at horizons of more than one or two years since the other dynamics die out quickly. Second, estimating the dominant root λ_s merely by plugging the ordinary least squares estimates of ρ_s and the ϕ_{si}s into (16) and solving for it yields a strongly biased estimate. For example, the median bias is .064 for $\lambda_s = .97$, $T = 60$, and $p = 0$; see Donald Andrews (1993, p. 148). One must therefore correct for the bias in some fashion.

Using local-to-unity asymptotic distribution theory, James Stock (1991) formulated a convenient method for doing so and for constructing confidence intervals. Using this method, I calculated unbiased estimates and 90 percent confidence intervals of the convergence rates for the per capita factor incomes of the contiguous US states. Figure 1 plots these convergence rates, which are arranged in ascending order.

[14] Carlino and Mills (1993) have presented evidence that the parameters in equation (5) may have shifted between the interwar and postwar period. Because parameter instability biases unit-root tests against

The estimated convergence rates average .1445 per year with a standard deviation of .1472 per year. The confidence intervals are wide, typically including both negative and very large positive values. Drawing strong inferences about the convergence rate of any given state is thus unwarranted. It is nonetheless reasonable to claim that about a quarter of the states have convergence rates of .2 or more. Furthermore, the uncertainty about the mean of these convergence rates is considerably less than the uncertainty about each state’s convergence rate. The central limit theorem suggests that the confidence interval for the mean of the estimated convergence rates should be about $1/\sqrt{48}$ times as wide as the average of the widths for the 48 individual states. The 90 percent confidence interval for the mean of the convergence rates is therefore approximately (.1248,.1660).

Using the cross-sectional approach described in section 2, Barro and Sala-i-Martin (1991, 1992, 1995) have estimated convergence rates for the US states of about two percent a year, much smaller than what is reported in the previous paragraph. I obtain similar estimates if I use their approach. For example, fitting equation (2) to the data and using the same *x*s as I employ in regression (21) below yields

(17) $$D\hat{y}_s = \underset{(.0175)}{.0386} - \underset{(.0008)}{.0117}\, y_s^0 + \underset{(.000588)}{.000600}\, HEART_s - \underset{(.00054)}{.00189}\, MOUNT_s$$

$$+ \underset{(.000090)}{.000438}\, COL_s, \quad R^2 = .9994, \quad SEE = .001305,$$

which implies a convergence rate of 2.32 percent a year.[15] I included the same *xs* in equation (17) as I found to be significant in equation (21) below.[16] My 1997b paper shows that if the parameters of equation (5) do not vary across states and if $p = 0$, then

(18) $$r = 1 - [1 + 68\,\text{plim}\,\hat{h} / (1 - w)]^{1/68},$$

where r is the true convergence rate, 68 is the number of years over which Dy is calculated, $\hat{h}$ is the ordinary least squares estimator of η, and $\omega \equiv \text{cov}(y^0,\mu|x)/\text{var}(y^0|x)$. A consistent estimate of ω for the data at hand is .2018.[17] Plugging this value and the

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rejection of the null hypothesis, the evidence is thus even more resoundingly against the null hypothesis than the analysis in the text would suggest.

[15] This figure is $1\text{-}(1\text{-}.0117\times 68)^{1/68}$, where 68 is the number of years that the growth rate *Dy* spans.

[16] *HEART* is one for states in what I call the *heartland* and zero for the other states; *MOUNT* is one for mountain states and zero for the other states; and *COL* is the average fraction of the population 25 years or older with at least four years of college in 1970, 1980, and 1990. The next subsection provides more detail.

[17] Squaring the standard error of estimate of the ordinary least squares regression of y^0 on an intercept, *HEART*, *MOUNT*, and *COL* produces a consistent estimate of $\text{var}(y^0|x)$. Multiplying the square of the

estimate of η in the regression (17) into equation (18) results in a convergence rate of 8.91 percent a year. Unfortunately, this estimate of the true convergence rate cannot be statistically distinguished from even 100 percent a year since plugging .2032 into equation (18) in lieu of .2018 would result in $r = 1$. (The difference between .2032 and .2018 is not statistically significant at any conventional significance level.) Therefore, cross-sectional regressions do not rule out even gigantic convergence rates.

4.3 What Affects Convergence Rates?

The characteristics of a state may affect its convergence rate. In particular, a state's convergence rate may depend on where it is and how well its population is educated. For example, Jess Benhabib and Mark Spiegel (1994) and Barro (1997) have argued that a more schooled population is more easily able to adopt new technologies from other regions and to transfer physical and human capital from low-return regions to high-return regions.

I therefore regressed the convergence rate on five regional dummy variables (HEART, SOUTH, PLAIN, MOUNT, and PACIF[18]), the average fractions of the population 25 and older with at least four years of high school (*HS*) and at least four years of college (*COL*) in 1970, 1980, and 1990, and the average number of years of schooling completed by this population (*SCH*) in 1940, 1950, and 1960. After dropping insignificant regressors, I obtained the regression

(19)
$$\hat{r}_s = \underset{(0.39)}{1.84} - \underset{(.083)}{.382}\ SOUTH_s - \underset{(.0064)}{.0312}\ HS_s + \underset{(.0108)}{.0314}\ COL_s,$$

$$R^2 = .6683, \quad SEE = .1551,$$

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standard error of estimate of equation (19) by the coefficient on μ in the regression of y^0 on an intercept, *HEART*, *MOUNT*, *COL*, and μ produces a consistent estimate of $\text{cov}(y^0,\mu|x)$. The arithmetic is $1.251(.1014/.2524)^2 = .2018$.

[18] *HEART* is one for Delaware, Illinois, Indiana, Maryland, Michigan, New Jersey, New York, Ohio, Pennsylvania, and Wisconsin and zero otherwise; *SOUTH* is one for Alabama, Arkansas, Florida, Georgia, Kentucky, Louisiana, Mississippi, North Carolina, Oklahoma, South Carolina, Tennessee, Texas, Virginia, and West Virginia and zero otherwise; *PLAIN* is one for Iowa, Kansas, Minnesota, Missouri, Nebraska, North Dakota, and South Dakota and zero otherwise; *MOUNT* is one for Arizona, Colorado, Idaho, Montana, Nevada, New Mexico, and Utah and zero otherwise; and *PACIF* is one for California, Oregon, and Washington. The excluded states are Connecticut, Maine, Massachusetts, New Hampshire, Rhode Island, and Vermont.

where $\hat{r}_s$ is the estimated convergence rate of state *s* and the figures in parentheses are heteroskedasticity-consistent standard errors. Not surprisingly, the estimates indicate that southern states tend to have lower convergence rates than other states. States in which large fractions of the population have four years of high school but not four years of college also tend to have low convergence rates. This result is inconsistent with those reported by Benhabib and Spiegel (1994) and Barro (1997) for samples of countries.

4.4 How Far Apart Are State Balanced Growth Paths?

The relative height of state *s*'s balanced growth path can be measured by the unconditional mean of $y_{st} - \bar{y}_t$. The standard deviation of these means across the US states is thus a reasonable measure of how far apart their balanced growth paths are. Equation (5) implies that μ_s, the unconditional mean of $y_{st} - \bar{y}_t$, takes the form

(20) $$\mu_s \equiv -\delta_s/\rho_s.$$

It can therefore be estimated by using ordinary least squares to fit equation (5) to the data for state *s* and plugging the resulting estimates of δ_s and ρ_s into equation (20). Carrying out these calculations yields estimated means with a cross-state standard deviation of .1580. Thus, the balanced growth paths for factor income are rather far apart.

The μ_ss may depend on the states' characteristics. I therefore regressed the μs on the same variables as I did for the convergence rates. After dropping insignificant regressors, I obtained the following:

(21) $$\hat{m}_s = \underset{(.083)}{-.593} + \underset{(.037)}{.135}\, HEART_s - \underset{(.0417)}{.0811}\, MOUNT_s + \underset{(.0055)}{.0381}\, COL_s,$$

$$R^2 = .6146, \quad SEE = .1014.$$

This regression indicates that per capita factor income tends to be higher in the heartland and lower in the mountain states than elsewhere. This finding may reflect the prevalence of manufacturing in the former and extractive industries in the latter. A college-educated population is also associated with higher per capita factor incomes. Each one percent of the population with four years of college adds about three percent. The size of the high-school population and years of schooling do not appear to have a separate effect, however. Interestingly, after adjusting for the size of the college-educated population, the South does not appear to have unusually low values of per capita factor income. Of course, in interpreting this regression, one should realize that the coefficient on *COL* could just as well reflect causation running from per capita factor income to schooling as the other way.

4.5 How Does the Cross-State Variance Evolve over Time?

Another way of investigating whether the per capita incomes of the contiguous US states converge toward parallel balanced growth paths is to examine how the cross-state variance of y_{st} evolves over time.[19] If $y_{st} - \bar{y}_t$ is covariance stationary for every state s,

$$V_t = \tfrac{1}{S}\sum_{s=1}^{S}(y_{st} - \bar{y}_t)^2 \qquad (22)$$

is also covariance stationary. In particular, it fluctuates around a constant positive mean. By contrast, under H_0, V is difference stationary with a positive drift rate; for proof, see my 1996 paper. The latter result is most easily seen when $y_{st} - \bar{y}_t$ is a driftless random walk in each region. In that case,

$$V_t = \tfrac{1}{S}\sum_{s=1}^{S}\sigma_s^2 + V_{t-1} + \tfrac{1}{S}\sum_{s=1}^{S}(u_{st}^2 - \sigma_s^2), \qquad (23)$$

where σ_s is the standard deviation of u_{st}. Putting these results together, one can assert that V has the representation[20]

$$\Delta V_t = \kappa + \gamma V_{t-1} + \sum_i \psi_i \Delta V_{t-i} + \varepsilon_t, \qquad (24)$$

where ε_t is an independently and identically distributed zero-mean error term with a positive finite variance and κ, γ, and the ψs are parameters such that $\kappa > 0$ and $-2 < \gamma \le 0$. The null hypothesis holds if $\gamma = 0$, and the alternative hypothesis holds if $\gamma < 0$. Note that V cannot fall indefinitely since it must be either stationary around a constant positive mean or nonstationary and upward trending.

[19] Danny Quah (1996) has used a related approach in which the cross-sectional distribution of per capita income is discretized and the intertemporal transitions between the points of the distribution are modeled as Markovian. This approach permits a richer description of how the cross-sectional distribution evolves over time than can be obtained from examining just the cross-sectional variance. Unfortunately, it presupposes covariance stationary, thereby ruling out the hypothesis testing carried out below.

[20] If the δs differ from zero as they would if the growth rates of states were endogenous, a linear time trend would also appear in equation (24). Its slope coefficient would be positive under the null hypothesis and zero under the alternative hypothesis. See my 1996 paper for proof.

Under the null hypothesis, the t-ratio for the least-squares estimator of γ converges in distribution to standard normal since $k > 0$. Furthermore, it diverges to $-\infty$ under the alternative hypothesis. Thus, a sufficiently negative t-ratio permits the null hypothesis to be rejected. For the sample sizes encountered in practice, however, the finite-sample distribution is intermediate between the Dickey-Fuller and standard normal distributions. As the ratio of κ to the standard deviation of ε_t rises from zero to infinity, the finite-sample distribution passes from the former to the latter. For example, the .05 critical values of the distribution with $T = 66$ are –2.913, -2.232, -1.882, -1.754, and -1.645 for ratios of 0, .4, 1.0, 2.5, and ∞. Therefore, if this ratio is known, Monte Carlo simulations can be used to obtain appropriate critical values.

Calculating an approximation to the ratio is straightforward. Suppose that u_{st} can be adequately represented as normal with identical variances σ^2 across the regions. In that case, $\kappa = \sigma^2$ and u_{st}^2 / s^2 is distributed as $\chi^2(1)$, which has a mean of one and a variance of two. Hence, given independence across the regions, the variance of ε_t is $2\sigma^4/S$, and the ratio in question is $\sqrt{S/2}$.

Figure 2 plots the cross-state variance of the logarithm of per capita factor income. Two features of the plot are noteworthy: innovations to the variance appear to dissipate quickly, and the variance has fallen a great deal over the sample period. The first observation suggests strong mean reversion, while the latter indicates that factor mobility has increased over time, thereby shifting the data generating process toward either faster convergence or less regional heterogeneity of immobile factors. If such changes have indeed taken place, assuming that equations (5) and (24) hold over the entire sample period with unchanged parameters is problematical. Such parameter instability decreases the likelihood of finding evidence of mean reversion, however. For this reason, the evidence for rapid convergence is even stronger than it would appear to be on first blush.[21]

[21] A fall in the variance resulting from structural shifts is not "convergence" in the sense that the term is used in this paper. Convergence implies that only about half the changes in V lower it unless initial conditions somehow place it far from its mean. *De novo*, it would be better to use "reversion" to refer to the tendency to approach balanced growth paths and "convergence" to refer to the effects of structural shifts like increasing mobility or decreasing transportation costs and trade barriers. The dead hand of past usage, however, makes this distinction more confusing than clarifying.

Pretesting revealed that one lag suffices to fit the data. The result from using ordinary least squares to fit equation (24) is

$$\Delta\hat{V}_t = \underset{(.00149)}{.00166} - \underset{(.0207)}{.0567}\, V_{t-1} + \underset{(.105)}{.278}\, \Delta V_{t-1}, \quad R^2 = .1881, \quad SEE = ..007407. \tag{25}$$

The *t*-ratio for the estimate of γ is –2.74. A Monte Carlo simulation with 100,000 iterations revealed that the .01 critical value of this test statistic is approximately –2.427. The evidence for convergence of the factor incomes of the contiguous US states to parallel growth paths is therefore strong. This finding confirms the other results of this section.

5. Comparison with Countries

It is instructive to compare the results reported in the previous section for US states with those that the literature has obtained for countries. As pointed out in section 2, there is little evidence for convergence across large groups of countries. If convergence is nonetheless accepted as given, my 1997b paper provides comparable estimates of convergence rates for 48 countries. The estimated convergence rates for most of these countries have wide confidence intervals, extending from somewhat negative values to substantial positive values. The mean of these estimates is 5.89 percent a year with a 90 percent confidence interval extending from 3.93 percent a year to 8.58 percent a year.[22] The convergence rates for the US states are therefore likely to be much larger than the convergence rates for these countries. A natural interpretation is that factor mobility is much higher across US states than it is across countries. In addition, the relative lack of barriers to trade in goods and the relatively low transportation costs within the US may play an important role. See the evidence reported in this volume by Antonio Fatas and Holger Wolf, which indicates that factors are much more mobile and trade barriers are much lower between US states than between countries.

Mean cross-country differences in log per capita income are much larger for countries than for US states. The standard deviation of the μs for the 48 countries in the sample for my 1997b paper is .9900, which is gigantic compared to .1580 for the per

[22] Indeed, the evidence that convergence occurs at all is largely confined to rich countries. See my 1996 and 1998 papers.

capita factor incomes of the contiguous US states. No doubt the low degree of international factor mobility and the high degree of interstate factor mobility in the US are key determinants of this difference.

In order to assess how schooling affects convergence rates and the levels of the balanced growth paths for per capita income, I regressed them on *SEC*, the average enrollment rate in secondary schools between 1950 and 1990.[23] The resulting regressions are

$$\hat{r}_c = \underset{(.0360)}{.0496} + \underset{(.0839)}{.0231}\, SEC_c \quad R^2 = .1387,, \quad SEE = .1506, \tag{26}$$

and

$$\hat{m}_c = \underset{(0.20)}{-1.23} + \underset{(0.42)}{3.19}\, SEC_c, \quad R^2 = .5569, \quad SEE = .6677. \tag{27}$$

Unlike for the states, the convergence rates appear to be insensitive to schooling. By contrast, schooling appears to be an important determinant of the μs for both countries and states. Interestingly, dummy variables for Latin America, Asia, and Africa, which by themselves are highly significant in a regression for $\hat{m}_c$, are completely insignificant after controlling for schooling.

6. Conclusions

The per capita factor incomes of the contiguous US states show a pronounced tendency to converge toward parallel balanced growth paths. Furthermore, the convergence is rapid on average, though the estimated convergence rates are widely dispersed across the states and quite imprecisely estimated for each individual state. Theory suggests that this rapid convergence results from the high factor mobility within the United States, though the absence of interstate barriers to trade in goods and low transportation costs may also play an important role. Notwithstanding the high factor mobility, differences in per capita income across the states are substantial. This result, however, is not surprising since the

[23] The countries included in the sample are Argentina, Australia, Austria, Belgium, Bolivia, Brazil, Canada, Chile, Columbia, Costa Rica, Denmark, the Dominican Republic, Ecuador, El Salvador, Finland, France, Germany, Greece, Guatemala, Honduras, India, Ireland, Italy, Japan, Kenya, Mexico, Morocco, the Netherlands, New Zealand, Nigeria, Norway, Pakistan, Panama, Paraguay, Peru, the Philippines, Portugal, Spain, South Africa, Sweden, Switzerland, Trinidad and Tobago, Thailand, Turkey, the United Kingdom, the United States, Uruguay, and Venezuela. My 1997b paper details the sources of the data on real GDP per worker and SEC_c.

economies of the states are quite heterogeneous. With such heterogeneity, per capita incomes would not be equalized even if all factor prices were equalized.

The per capita incomes of broad groups of countries show no pronounced tendency to converge toward balanced growth paths. To the extent that convergence does takes place, it is fairly slow on average. As was true for the states, the estimated convergence rates are widely dispersed across the countries and estimated with considerable imprecision for each individual country. Theory suggests that this fairly slow, or even nonexistent, convergence results from the virtual absence of labor mobility across national frontiers and perhaps from low capital mobility, barriers to trade in goods, and significant transportation costs. Finally, differences in labor productivity across countries are gigantic, greatly exceeding those across the contiguous US states. No doubt the absence of labor mobility accounts for the lion's share of these differences, though the heterogeneity of the economies of the countries probably also exceeds that of the states.

These results suggest that the movement toward unified markets in goods and factors in Europe should lead the per capita incomes of the individual countries to converge toward parallel balanced growth paths if they do not already do so. Moreover, to the extent that they do, the rates at which the convergence takes place should increase, the balanced growth paths should be pulled toward each other, and the cross-sectional variance of the logarithms of per capita income should fall.

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Table 1. Real Per Capita GDP for Country Groups in 1990 Geary-Khamis Dollar

Country Group	***1820***	***1870***	***1913***	***1950***	***1973***	***1992***
Western Europe	1292	2110	3704	5123	12288	17384
Western Offshoots	1205	2440	5237	9255	16075	20850
Southern Europe	806	1111	1753	2025	6029	8273
Eastern Europe	750	1030	1557	2604	5742	4608
Latin America	715	800	1515	2614	4750	5294
Asia	550	580	742	727	1680	3239
Africa	450	480	575	792	1274	1318

Source. Table E-3 of Maddison (1995).

Table 2. Growth Rates of Real Per Capita GDP for Country Groups in Percent per Year

Country Group	1820-1870	1870-1913	1913-1950	1950-1973	1973-1992
Western Europe	0.98	1.31	0.88	3.80	1.83
Western Offshoots	1.41	1.78	1.54	2.40	1.37
Southern Europe	0.64	1.06	0.39	4.74	1.67
Eastern Europe	0.79	0.97	1.38	3.44	-1.16
Latin America	0.22	1.49	1.47	2.60	0.57
Asia	0.11	0.57	-0.05	3.64	3.46
Africa	0.13	0.42	0.87	2.07	0.18

Source. Derived from data in Table 1 above.

Table 3. Summary Statistics on the *t*-Ratios for the Contiguous U.S. States

Statistic	*Sample*	*Dickey-Fuller Distribution*
Mean	-2.250	-1.528
Standard Deviation	0.837	0.866
5th Percentile	-3.994	-2.911
10th Percentile	-3.779	-2.597
25th Percentile	-2.878	-2.088
50th Percentile	-2.286	-1.554
75th Percentile	-1.702	-0.992
90th Percentile	-1.275	-0.413
95th Percentile	-0.903	-0.042

Notes. A common lag length (p) of three years was chosen for each state. Pretesting indicated that this choice is adequate. Estimation was by ordinary least squares over the sample period 1933-1996. A Monte-Carlo simulation with 1,000,000 iterations estimated the values reported in the last column.

Figure 1. Median Estimates and 90% Confidence Intervals
of Convergence Rates for Factor Income

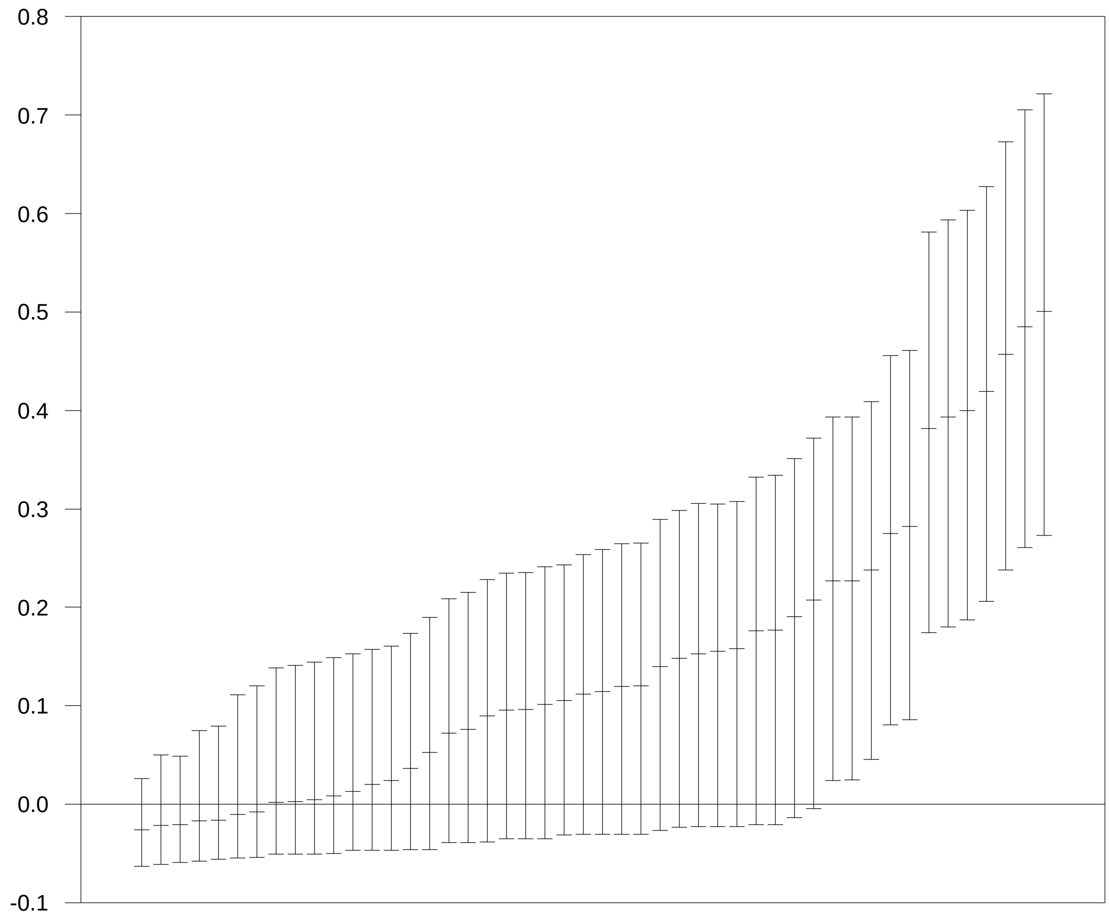

Figure 2. Cross-State Variance
of Log Per Capita Factor Income, 1929-1996