PN-ABD-350

Economic Growth and Convergence across the United States

Robert J. Barro and Xavier Sala i Martin

Harvard University

September 1989

Preliminary

1

Introduction

Forthcoming.

A Model of Growth and Convergence

Two concepts of convergence appear in discussions of economic growth across countries or regions. In one view convergence applies if a poor country tends to grow faster than a rich one, so that—other things equal—the poor country tends to catch up with the rich one in terms of the level of per capita income. The second concept concerns the cross-sectional dispersion of per capita income. In this context, convergence occurs if the dispersion—measured say by the standard deviation of the logarithm of per capita income across a group of countries—declines over time. Convergence of the first kind (poor countries tending to growth faster than rich countries) works toward convergence of the second kind (reduced dispersion of incomes), but is offset by new disturbances that tend to increase dispersion.

To illustrate within a simple log-linear model, suppose that the growth rate of real per capita income for country or region i in period t is given by

$$(1) \qquad \log(y_t^i/y_{t-1}^i) = \alpha - \beta \cdot \log(y_{t-1}^i) + u_t^i$$

where y_t^i is the level of real per capita income or product. The random variable u_t^i has mean zero, variance σ_{ut}^2, and is distributed independently of $\log(y_{t-1}^i)$.

Convergence in the first sense applies if $\beta > 0$—we refer to this condition henceforth as β-type convergence. (We also assume $\beta < 1$, which means that convergence is not so strong to eliminate the positive serial correlation in $\log(y_t^i)$.) β-type convergence applies in the standard neoclassical growth model (Solow, 1956, Koopmans, 1965, Cass, 1965) if we think of countries as closed economies that differ only by their starting capital-labor ratios, k_t^i (see King and Rebelo, 1989, for a discussion). Because of diminishing returns to capital, countries with higher k_t^i have lower marginal products of capital and thus tend to have less investment and lower per capita growth rates. Abstracting from the stochastic term u_t^i, all countries would tend asymptotically to the same steady-state values of capital and output per worker.[1] (The condition $\beta < 1$ rules out oscillations about the steady state.) With open economies, the convergence coefficient β would tend to be larger because capital would migrate toward places where its marginal product is higher (i.e. toward the poor countries), while labor would move toward places where its marginal product is higher (i.e. toward the rich countries).

Let σ_t^2 be the variance of $\log(y_t^i)$ for date t. Equation (1) implies that σ_t^2 evolves over time in accordance with the first-difference equation,

---

[1] Equation (1) implies, if $u_t^i = 0$ for all t, that the per capita growth rate approaches zero aymptotically. The specification could be modified to allow for nonzero per capita growth in the steady state. For example, the term $\beta \cdot \log(y_{t-1}^i)$ could be replaced by $\phi(y_{t-1}^i)$, where $\phi'>0$, but where $\phi(y_{t-1}^i)$ is bounded from above as y_{t-1}^i approaches infinity. This revised formulation could accommodate nonzero per capita growth in the steady state.

$$\sigma_t^2 = (1-\beta)^2 \cdot \sigma_{t-1}^2 + \sigma_{ut}^2. \tag{2}$$

If σ_{ut}^2 is constant, the cross-sectional variance of income approaches a constant. In particular, letting σ_0^2 denote the variance of $\log(y_0^i)$ at some initial date 0, the solution of the first-difference equation (2) is

$$\sigma_t^2 = \frac{\sigma_u^2}{1-(1-\beta)^2} + [\sigma_0^2 - \frac{\sigma_u^2}{1-(1-\beta)^2}] \cdot (1-\beta)^{2t} \tag{3}$$

where σ_u^2 is the constant variance of the shocks. We assume that the sample size is large enough so that the measured cross-sectional variance of $\log(y_t^i)$ corresponds to the population variance σ_t^2. Then, if $0 < \beta < 1$, as we assume, the cross-sectional variance approaches the steady-state value, $\bar{\sigma}^2 = \sigma_u^2/[1-(1-\beta)^2]$, which rises with σ_u^2 but declines with β. That is, a higher convergence coefficient β reduces $\bar{\sigma}^2$, but the value of $\bar{\sigma}^2$ depends also on the variance of the shocks, σ_u^2. The variance σ_t^2 falls (rises) over time if σ_0^2 is greater than (less than) the steady-state value, $\bar{\sigma}^2$. Thus, a positive coefficient β (β-type convergence) does not necessarily imply a falling σ_t^2 (which we refer to as σ-type convergence).

Figure 1 shows the behavior of σ_t^2 over time for σ_0^2 above or below $\bar{\sigma}^2$. The convergence coefficient used, $\beta = .03$ per year, corresponds to estimates discussed later for the United States. With $\beta = .03$, it takes about 11 years to eliminate 50% of the initial gap between σ_0^2 and $\bar{\sigma}^2$, and about 23 years to eliminate 75% of the gap.

The variance of the shocks, σ_{ut}^2, may itself be a random variable. For example, σ_{ut}^2 may rise temporarily above its mean because of an aggregate disturbance that affects countries or regions differentially. In earlier times where agriculture was a major component of U.S. GNP, this shift could reflect a harvest failure. More recently, the sharp movements in the relative price of energy have similar effects.

Suppose that the cross-sectional variance of income, σ_t^2, is initially at the steady-state value, $\bar{\sigma}^2$, corresponding to the constant shock variance σ_u^2. Then a temporary increase in σ_{ut}^2 above σ_u^2 generates an interval of gradual rise in σ_t^2. If σ_{ut}^2 returns to the value σ_u^2, σ_t^2 declines gradually toward its original steady-state value. These kinds of movements in σ_t^2 occur even if the convergence coefficient, β, is constant. However, an important assumption is that the individual shocks, u_t^i, are independent of $\log(y_{t-1}^i)$. In the case where the increase in σ_{ut}^2 reflects a harvest failure or an oil shock or other disturbances that affect regions differentially, u_t^i may be correlated with y_{t-1}^i (for example, if the oil-producing regions have above-average per capita income). In this case the estimated β-coefficient (but not the true coefficient) is affected by this correlation. We consider this possibility below in our interpretation of results across the U.S. states.[2]

Convergence across the U.S. States

We now use the simple model described above to assess the behavior of economic growth and convergence across the U.S. states. For this purpose we have available two concepts of state per capita income or product. The first

[2]It is unfortunate that the term U.S. states equals United States states. But United States by itself has a different meaning and American states is clearly unsatisfactory. Confederate states would have worked okay.

measure is per capita, nominal state personal income. These data are available for the 48 continental U.S. states from 1929 to 1988 (although we presently lack figures for 1930-32, 1934-38, and 1941-42). We can deflate the nominal values by the U.S. deflator for national income for each year.

The second concept—available from 1963 to 1986—is nominal per capita gross state product, GSP. This variable, which is analogous to gross domestic product, measures production within each state. We can deflate the nominal figures by the aggregate gross state product deflator for each year.

Note that relative prices or price indices for individual states are unavailable as deflators for state income or product. Thus, comparisons of levels of measured real income or product across the states depend on conditions of purchasing-power parity. For comparisons of growth rates, it is only necessary that purchasing-power parity hold in a relative sense.

The main differences between the two concepts of state per capita income involve transfer payments and capital income. Transfers (within the state or from the federal government) appear in personal income but not in GSP. Personal income includes corporate net income only when individuals receive payment as dividends, whereas GSP encompasses corporate profits and depreciation. (Neither concept includes capital gains.) In addition, GSP attributes capital income to the state where the business activity occurs, whereas personal income attributes it to the state of the asset holder. Some of these locational differences apply also to labor income, although—except for a few cities—the locations of the business and worker are likely to be in the same state.

Table 1 shows cross-state regressions of the form of equation (1). We consider initially the presence of β-type convergence across the states in the long run. For the longest possible sample, 1929-88 for personal income, the presence of convergence—in the sense of a negative coefficient on the log of initial income—is dramatic. Using the annual average growth rate of state per capita real personal income over 59 years as the dependent variable, the estimated coefficient of $\log(y_{1929})$ on line 1 is -.0107, s.e. = .0007. Figure 2 shows this relationship graphically.

Convergence also appears using real per capita gross state product, GSP, over the full sample that is available, 1963-86. Using the annual average growth rate of GSP over 23 years as the dependent variable, the estimated coefficient of $\log(GSP_{1963})$ on line 2 of the table is -.0172, s.e. = .0032 (see Figure 3).[3]

We now consider whether β-type convergence shows up over shorter periods. Lines 6, 8, 10, and 12 of Table 1 show the convergence coefficients for GSP growth over five-year intervals: 1965-70, 1970-75, 1975-80, and 1980-85. The estimated coefficients appear to be unstable; the values for the respective intervals are -.034 (s.e. = .008), -.036 (.016), .017 (.014), and -.054 (.013).[4] A test of the hypothesis that the four coefficients are equal (allowing for separate constant terms in each time period) leads to the

[3]For real personal income, the estimated coefficient for the 1963-86 sample is -.0118, s.e. = .0029.

[4]Each of these coefficients refers to the relation between the log of initial real per capita product and the average growth rate of real per capita product over the subsequent five years. Thus, it is plausible that the coefficients would be equal. With different averaging intervals—such as the 59 years from 1929 to 1988 versus any of the 5-year periods—the theoretical equation (1) implies that the coefficients would decline as the interval gets longer.

statistic, $F^3_{184} = 5.1$, which is significant at the 1% level and therefore leads to rejection of the null hypothesis.

Notice that the estimated coefficient on $\log(GSP_0)$ for 1975-80 (line 10) has the wrong (positive) sign. We conjectured that this result related to the sharp increase in the relative price of energy in 1979-80. In particular, states with substantial production in energy became relatively high in per capita state product after the 1973 oil crisis. Thus, the shock in 1979-80 would not only increase the dispersion of GSP, it would also bias the estimated coefficient of $\log(GSP_{1975})$ in the regression for the growth rate over 1975-80 in a positive direction. Since the reduction in the relative price of oil in the 1980s was bad for the high-income oil states, the estimated coefficient of $\log(GSP_{1980})$ for the 1980-85 regression would be biased downward. Hence this mechanism could also explain why the estimated coefficient on $\log(GSP_{1980})$ for 1980-85 (line 12) is more negative than those for 1965-70 and 1970-75.[5]

Formally, let S_t be a random variable that represents an economy-wide disturbance for period t. For example, S_t could reflect the relative price of oil as determined in world markets. Then equation (1) could be modified to

$$\log(y^i_t/y^i_{t-1}) = \alpha - \beta\cdot\log(y^i_{t-1}) + \gamma^i\cdot S_t + u^i_t \tag{4}$$

[5]This effect is even more important for a regression over 1981-86, which covers the sharp decrease in the relative price of oil in 1986. Over this sample, the estimated coefficient of $\log(GSP_{1981})$ is -.0893, s.e. = .0164.

where γ^i measures the effect of the aggregate disturbance on the growth rate of GSP for state i. For example, if a positive value of S_t signifies an increase in the relative price of oil, γ^i would be positive for states that produce a lot of oil. (The coefficient γ^i would tend to be negative for states that produce goods, such as automobiles, that use oil as an input.) We think of the coefficients γ^i as distributed over the states with constant mean $\bar{\gamma}$ and constant variance σ_γ^2.

Equation (4) implies

$$(5) \qquad \sigma_t^2 = (1-\beta)^2\sigma_{t-1}^2 + \sigma_u^2 + S_t^2\cdot\sigma_\gamma^2 + 2S_t(1-\beta)\cdot COV[\log(y_{t-1}^i), \gamma^i]$$

where the variances and covariance are conditioned on the current and past realizations of the aggregate shocks, S_t, S_{t-1}, ... If $COV[\log(y_{t-1}^i), \gamma^i]$ were zero, equation (5) would correspond effectively to equation (2). Shifts in σ_{ut}^2 would result from realizations of S_t that were larger or smaller in magnitude. It also follows in this case that estimates of the coefficient β in equation (4) would not be systematically related to the realizations of S_t. That is, the composite error term, $\gamma^i\cdot S_t + u_t^i$, would be uncorrelated with the regressor, $\log(y_{t-1}^i)$, and an OLS estimate of the coefficient β would be consistent.

Suppose, alternatively, that $COV[\log(y_{t-1}^i), \gamma^i] > 0$. That is, if a positive S_t represents an increase in the relative price of oil, states that produce a lot of oil ($\gamma^i > \bar{\gamma}$) have above average per capita product in period t-1.[6] In that case the estimated coefficient on $\log(y_{t-1}^i)$ in equation (4)

[6]The share of production of crude oil and natural gas in gross state product is little correlated with per capita product from 1963 to 1973. Then, with the oil shocks, the correlation rises from 0.1 in 1973 to 0.4 in 1975 and 0.7

would be positively or negatively biased as S_t is positive or negative. That is, we would tend to underestimate β-type convergence in a period where the oil price rises ($S_t > 0$) and vice versa.

If $COV[\log(y^i_{t-1}), \gamma^i] > 0$, the derivative of the covariance term in equation (5) with respect to S_t is positive. Therefore, the derivative of σ^2_t is positive if $S_t > 0$, but may be positive or negative if $S_t < 0$. If $\log(y^i_{t-1})$ and γ^i were perfectly correlated—so that $COV[\log(y^i_{t-1}), \gamma^i] = \sigma_{t-1} \cdot \sigma_\gamma$—$\sigma^2_t$ would be monotonically increasing in S_t. Thus, if oil-producing states tend to have above-average per capita product, an increase in the relative price of oil has a positive effect on the cross-sectional dispersion of per capita product. A decrease in the relative price of oil may lower this dispersion, and surely lowers it if the correlation between $\log(y^i_{t-1})$ and γ^i is unity.

Figure 4 and Table 2, column 1 show how the (unweighted) cross-sectional standard deviation of log(GSP), σ_t, evolved from 1963 to 1986. Note that σ_t declined steadily from 1963 to 1972. This σ-type convergence could reflect the dynamics worked out in equation (3) and Figure 1 if σ_t began above its steady-state value. Figure 5 and Table 2, column 2 show the behavior of σ_t based on state personal income. In this case σ_t declined dramatically during World War II, and fell gradually from 1950 until the mid 1970s.

Figure 3 indicates that σ_t for log(GSP) rose from 1972 to 1975, increased sharply from 1978 to 1981, and fell rapidly from 1981 to 1986.[7] As discussed

in 1981. With the subsequent decline in oil prices, the correlation falls after 1981 to reach 0.5 in 1984 and 0.1 in 1986.

[7]The pattern for state personal income, shown in Figure 5, is different here—σ_t reaches its trough in 1976 and then rises steadily through 1988. The differing behavior from GSP does not relate to government transfers. If total transfer receipts from government are deducted from state personal

before, the increases in oil prices could generate the increase in σ_t up to 1981. Once oil prices stopped rising, the theory predicts that σ_t would fall back toward its steady-state value. However, the decline in oil prices likely accelerated the reduction in σ_t that shows up after 1981 in Figure 4. We get this effect in the theory if $\log(y^i_{t-1})$ and γ^i are highly positively correlated (see n. 6 above).

To isolate the pure β-type convergence related to the level of initial per capita product, we want to hold fixed the effects—such as those related to oil shocks—that show up in the term, $\gamma^i \cdot S_t$, in equation (4). We proceed by constructing a variable, called STRUCTURE, to hold constant the effects of each state's structural composition of output. We broke down each state's GSP for each year into the shares contributed by 54 sectors. Then we multiplied the shares of each sector in state i for year t-1 by the national growth rate of the sector from year t-1 to year t. By adding up these multiples we determined what state i's growth rate of GSP would have been from year t-1 to year t if each of the state's sectors had grown at the same rate as the national average. In particular, the variable STRUCTURE holds constant the positive (negative) effect of higher (lower) oil prices on the GSP growth of oil producers. That is, it would pick up the term $\gamma^i \cdot S_t$ in equation (4).

---

income, the pattern for σ_t remains similar to that shown in Figure 5. Thus, it seems that the divergent behavior would relate to capital income, which enters differently into personal income and GSP.

Lines 7, 9, 11, and 13 of Table 1 show that the coefficients of $\log(GSP_0)$ are more stable over sub-periods if the variable STRUCTURE is held constant. For the samples, 1965-70, 1970-75, 1975-80, 1980-85, the respective estimated coefficients are now -.027 (.007), -.036 (.012), -.024 (.008), and -.036 (.010). An F-test for equality of these coefficients (allowing each sub-period to have individual constants and coefficients on STRUCTURE) is $F^3_{180} = 0.3$, which is not significant. The pooled estimate of the convergence coefficient from the four 5-year samples (line 16) is -.0312, s.e. = .0046. (With a single coefficient on STRUCTURE—which would not be rejected by the data—the estimate on line 15 is -.0294, s.e. = .0044.)

As discussed before, the strong impact on estimated convergence coefficients arises when the structural effects are correlated with initial GSP. Thus, the effects for 1975-80 and 1980-85 arise because oil producers have high starting GSP in each sub-sample. However, the signs of the effects are opposite in the two periods since oil prices rose in the first and fell in the second. Over long periods, the effects of STRUCTURE on growth tend to be less correlated with initial GSP. Hence the estimated convergence coefficients are less sensitive to the inclusion of the structural variable. For example, with the variable STRUCTURE omitted, the pooled estimate for the coefficient of $\log(GSP_0)$ over the four 5-year intervals is -.0295, which is virtually identical to that obtained with STRUCTURE held constant.[8] Similarly, over the 1963-86 period, the estimated coefficient of $\log(GSP_{1963})$

[8] However, as noted, we would reject the hypothesis that the coefficient of $\log(GSP_0)$ was stable over the sub-periods if the variable STRUCTURE were excluded.

with STRUCTURE included (-.0205 in line 3) does not differ greatly from that with STRUCTURE omitted (-.0172 in line 2).

We mentioned that one mechanism for convergence in state per capita product would be the migration of labor toward the states with high GSP. We computed annual population growth rates due to net migration for each state. For a pooled sample of the five-year intervals from 1965-70 to 1980-85 (with individual constant terms), the estimated regression coefficient of this migration variable on $\log(GSP_0)$ is .0101, s.e. = .0043. Hence the estimated coefficient is significantly positive, indicating net migration toward the states with high initial GSP. We plan to explore these effects more fully along with an analysis of overall population growth across the U.S. states.

Convergence across 114 Countries

There are interesting parallels and differences between the convergence patterns across the U.S. states and those across countries. It is well known that growth rates of real per capita GDP are uncorrelated with the starting level of real per capita GDP across a large number of countries in the post-World War II period. Figure 6 uses the Summers-Heston (1988) data on real per capita GDP to show this pattern for 114 countries from 1960 to 1985. There is, if anything, a positive association between the per capita growth rate and the starting level of per capita GDP.[9] Table 3, line 1 indicates

---

[9]Typical versions of this diagram, such as Romer (1989, p. 64), use the level of real per capita GDP rather than the logarithm on the horizontal axis. When the logarithm is used, the diagram no longer has the dramatic triangle shape where the growth rates appear to be much more spread out at the left end. Put another way, although the high GDP countries on the right of Figure 6 have a lower variance of growth rates, the range of these countries on the right occupies a much smaller fraction of the horizontal axis than it would if the level of GDP were measured on the axis.

that the regression coefficient on $\log(GDP_{1960})$ is .005, s.e. = .002. (The dependent variable here is the average growth rate of per capita GDP from 1960 to 1985.)

A previous study (Barro, 1989b) found that the estimated coefficient on 1960 GDP became significantly negative if measures of initial human capital were held constant. Table 3, line 2 adds the 1960 values of primary and secondary school-enrollment rates as proxies for the starting amount of human capital. With these variables included, the estimated coefficient on $\log(GDP_{1960})$ is -.0090, s.e. = .0028.[10] Thus, the magnitude of the estimated convergence coefficient is much smaller than that from a regression over 1963-86 for the U.S. states (.0172 in line 1 and .0205 in line 2 of Table 1). Given the greater mobility of factors across the U.S. states than across the 114 countries, this result seems reasonable.

The role of the human capital variables can be understood by modifying the growth equation (1) (now applicable to countries) to

$$\log(y_t^i/y_{t-1}^i) = \alpha - \beta\cdot\log(y_t^i) + \delta\cdot h_t^i + u_t^i \tag{6}$$

where h_t^i is a measure of human capital for country i (measured say by school-enrollment rates) and $\delta > 0$. Thus equation (6) allows the growth rate of per capita GDP to respond positively to the level of human capital. For

[10]The results in Barro (1989b) use the level of GDP_{1960} instead of the logarithm, and also include additional explanatory variables. These changes improve the fit of the regression, but the results on convergence are not greatly affected.

theoretical discussions of this type of effect, see Becker and Murphy (1990) and Barro (1989a). The variables h_t^i and $\log(y_{t-1}^i)$ are positively correlated across countries. Therefore, if h_t^i is omitted from a cross-country regression, the estimated coefficient on $\log(y_{t-1}^i)$ is positively biased. When proxies for h_t^i are included, the estimated coefficient on the logarithm of initial GDP tends to be negative. However, the interpretation here is that a poor country tends to grow faster than a rich country if the two countries have the same level of human capital (or, more generally, if the poor country is abundant in human capital in relation to its level of per capita GDP). If the two countries have the quantities of human capital that correspond typically to the respective levels of per capita GDP, the per capita growth rates tend to be about the same, as suggested in Figure 6.

Table 3 shows also cross-country regressions over 5-year intervals from 1960-65 to 1980-85. With school-enrollment rates included, the hypothesis of stability for the coefficients of $\log(GDP_0)$ is rejected at the 1% level: the test statistic is $F_{550}^4 = 3.2$. (These results allow each sub-period to have separate coefficients on school-enrollment rates and constant terms.)

In the case of the U.S. states, we accounted for the instability of estimated convergence coefficients over 5-year periods by arguing that an omitted factor—energy shocks—was first, more important in some periods than others, second, it exerted differential effects on state growth rates, and third, it was correlated with the initial level of per capita product but in a way that varied over time. Looking at the cross-country estimates for the coefficients of $\log(GDP_0)$ on lines 4, 6, 8, 10, and 12 of Table 3, the outliers are the positive coefficient (.0082) for 1960-65 and the sharply negative coefficient for 1980-85 (-.0166). It does not appear that energy shocks would account for this pattern, although it is possible that other structural influences would matter for the countries. At this point we lack a variable analogous to STRUCTURE that we can use for the cross-country sample. However, our cross-country results suggest that the omission of a variable like STRUCTURE may not matter much for the long-term results.

It is possible that other omitted factors would affect cross-country growth in a way analogous to that for the human capital variable in equation (6). For example, Greenwood and Jovanovic (1989) have stressed the role of financial intermediation in fostering economic growth. As with human capital, financial development is likely to have a positive effect on growth and is also positively correlated with initial per capita product. Thus, the omission of this variable tends to bias the estimated coefficient on $\log(GDP_0)$ in a positive direction. As with human capital, the concept of convergence here is that a country with low starting income tends to grow fast if it is abundant in financial intermediation in relation to its initial level of income. If a country has the amount of financial development (and human capital) that is typical for its level of income, the low-income country would not tend to growth especially fast (as Figure 6 suggests).

Table 2, column 3 shows that the cross-sectional (unweighted) standard deviation of the logarithm of GDP for the 114 countries increased steadily from 1960 to 1985. Column 4 of the table shows that the standard deviation also increased slowly from 1950 to 1960 for the 60 countries for which the pre-1960 data are available. Not surprisingly, the level of these standard deviations is far greater than that across the U.S. states. (The concepts are, however, not strictly comparable because the Summers-Heston numbers deflate each country's GDP by a PPP-based price deflator, while the cross-state numbers divide the nominal figures by a national price index.)

Convergence across OECD Countries

The β-type convergence across the U.S. states was much more pronounced than that across the 114 countries. We are inclined to explain this difference by greater factor mobility across the states, which we think relates to a common central government, language, and other elements and secondarily to geographical proximity. To check these ideas, we considered the OECD countries as a sample that would likely be intermediate between these two cases. Unlike the U.S. states, the OECD countries lack a common central government, language, etc. But this group of countries would be more homogeneous in many respects than the 114 countries that we just considered.

Table 4 shows results for the 20 original OECD countries (the current membership except for Australia, Finland, Japan, and New Zealand).[11] Over the period 1960-85 with the 25-year annual average of per capita GDP growth as the dependent variable, the estimated coefficient of $\log(GDP_{1960})$ is -.0085, s.e. = .0022 (line 1). Figure 7 shows this relation graphically. Unlike the sample of 114 countries, but like the sample of U.S. states, there is evidence for β-type convergence across the OECD countries even with no other variables held constant.

[11] Romer (1989, p. 66) and DeLong (1988) have argued that coefficients on initial GDP in growth regressions are biased downward if one selects a sample based on GDP being high in the terminal period. The sample of 20 original OECD countries is based on membership determined as of 1960, that is, at the beginning of the sample. Thus, we avoid the possibility that the four countries admitted later to the OECD were, in some sense, chosen because of high GDP later on. (This possibility seems pertinent mainly for Japan.)

With the 1960 values of primary and secondary school-enrollment rates included (Table 4, line 2), the coefficient of $\log(GDP_{1960})$ rises in magnitude to -.0124, s.e. = .0028. These results are intermediate between those for the 114 countries (the comparable coefficient for 1960-85 in Table 3, line 2 is -.0090) and those for the 48 U.S. states (the coefficient for 1963-86 in Table 1, line 2 is -.0172). As suggested before, a possible interpretation is that the mobility of factors across the OECD countries is intermediate between that for the 114 countries and that for the U.S. states.

Table 4 also shows results for the OECD countries over 5-year intervals from 1960-65 to 1980-85. As with the U.S. states, the estimated coefficient on $\log(GDP_0)$ (with no other variables held constant) is negative through the mid 1970s, but becomes positive for 1975-80. Unlike the U.S. states, the point estimate for the OECD countries remains positive for 1980-85. With the school-enrollment variables included, a test of equality for the coefficients on $\log(GDP_0)$ over the 5-year sub-periods for the OECD countries leads to the statistic $F^4_{80} = 3.2$. (The estimation allows each sub-period to have separate coefficients on the school-enrollment variables and the constant terms.) Thus, the hypothesis of equality is rejected at the 1% level. This result may reflect the differential impact of energy shocks or other aggregate disturbances, which were held constant for the U.S. states, but not across the OECD countries.

Table 2, column 5 shows the behavior of the (unweighted) cross-sectional standard deviation of log(GDP), σ_t, for the 20 OECD countries. As with the U.S. states, σ_t declines steadily until the mid 1970s. Unlike the U.S. states, σ_t for the OECD remained roughly constant from 1975 to 1985. The difference may reflect the greater differential impact of energy shocks on the U.S. states than on the OECD. The level of σ_t for the OECD is, not surprisingly, intermediate between that for the U.S. states and that for the 114 countries.

Human Capital for the U.S. States

We have not had great success in isolating influences of human capital for the U.S. states that parallel those for the 114 countries or the OECD countries. We have collected U.S. Census data on fractions of the adult population (25 years and older) in each state that attained various levels of education by 1960. These variables are proxies for initial human capital in each state.

Holding fixed $\log(GSP_{1963})$, the variable COLLEGE (fraction of the adult population with a completed college education in 1960) is significantly positive (.107, s.e. = .044) for the average growth rate of GSP from 1963 to 1986—see Table 5, column 1. Because of the positive correlation between COLLEGE and $\log(GSP_{1963})$, the inclusion of COLLEGE raises the magnitude of the estimated coefficient on $\log(GSP_{1963})$ from .0172 (line 2 of Table 1) to .0224.

Holding fixed the variable COLLEGE, other measures of educational attainment—fractions of the population in 1960 with some college, completed secondary school, etc.—were insignificant for growth. Also, if the variable STRUCTURE is added (column 2 of Table 5), the estimated coefficient of COLLEGE is no longer significantly different from zero (.056, s.e. = .051). The results are similar in this respect for the 5-year sub-periods or for a pooled sample of the 5-year periods. The variable COLLEGE tends to be positive if STRUCTURE is omitted, but insignificant when STRUCTURE is included. This result was unexpected, since it was not obvious that college attainment and the structural growth variable would be positively correlated. (The correlation of COLLEGE with STRUCTURE measured over 1963-85 is .35.) It is possible that similar results would show up for the 114 countries or the OECD sample, but we do not have the variable STRUCTURE in these cases.

We plan to compile the necessary data to see whether school-enrollment rate variables—that is, the human capital measures used across countries—work differently from the educational attainment variables for the U.S. states. However, the attainment variables seemed, a priori, to be better proxies than school-enrollment rates for initial stocks of human capital.

Effects of Government Expenditure

Barro (1989b, Table 1) reports a number of results concerning the interplay between government and growth across countries. For 98 countries with available data, the estimated regression coefficient for the growth rate of per capita GDP (1960-85) on the ratio of government consumption to GDP (averaged over 1970-85) was -.12, s.e. = .03.[12] Variables that reflected political instability (numbers of revolutions, coups, and per capita assassinations) were significantly negative, as was a proxy for distortions of market prices (based on the Summers-Heston (1988) deflator for investment goods in 1960). These kinds of effects can be interpreted within the model of Barro (1990), where taxation, threats on property rights, and other

---

[12]However, for the 20 OECD countries the estimated coefficient on the government consumption variable differs insignificantly from zero: .012, s.e. = .028.

distortions can have a negative effect on the long-run growth rate. Unfortunately, some of the results also have a different interpretation, such as the reverse effect of economic growth on political stability.

For 76 countries with available data, per capita growth was insignificantly related to the ratio of public to total investment (Barro, 1989b, Table 4). However, this result does not necessarily imply that public investment is unimportant for growth. Barro (1990) shows that if governments choose productive spending to maximize the economy's per capita growth rate, the correlation between the growth rate and the share of output devoted to this spending would be close to zero.

Table 5 shows some preliminary results related to government spending across the U.S. states. The dependent variable in the regressions is the growth rate of per capita gross state product from 1963 to 1986. We have a proxy for initial public capital stock in each state, based on public investment from 1950 to 1962. The variable PUBI is the average from 1950 to 1962 of the ratio of state capital expenditures to state personal income.[13] (Because the data on capital expenditures by local governments are unavailable, the figures refer only to state governments.) The estimated coefficient of PUBI in column 3 of the table is .087, s.e. = .086. The dominant component of PUBI is highway expenditures, and the results are similar if only capital expenditures on highways are used.

The variable GOV is the average from 1963 to 1986 of the ratio of total expenditures by state and local governments in each state (netting out

[13]We cannot divide by gross state product because data on this variable are unavailable before 1963. We also lack information on private investment by state, except for investment expenditures in manufacturing.

transfers from state to local governments) to gross state product. The estimated coefficient of GOV in Table 5, column 4 is -.080, s.e. = .031. With GOV held constant, the estimated coefficient of PUBI rises to .14, s.e. = .08.

We have available the breakdown of state government expenditures, and to a lesser extent local government expenditures, into various functional components. For example, we have figures on spending for education, welfare, highways, police, and so on. We are presently studying the relation of spending in various categories to state economic growth. The key problem in this analysis is to hold fixed the endogenous parts of government spending. For example, police spending responds to the amount of crime, highway spending to the number of mountains and to prospects for economic growth, and so on. Thus, it is difficult to isolate the independent influence of government activities on economic growth. We do not claim to have solved this problem.

Conclusions

Forthcoming.

References

Barro, R.J., "A Cross-Country Study of Growth, Saving, and Government," National Bureau of Economic Research, working paper no. 2855, February 1989a.

---------- "Economic Growth in a Cross Section of Countries," unpublished, Harvard University, August 1989b.

---------- "Government Spending in a Simple Model of Endogenous Growth," *Journal of Political Economy*, forthcoming, 1990.

Becker, G.S. and K.M. Murphy, "Economic Growth, Human Capital, and Population Growth," *Journal of Political Economy*, forthcoming, 1990.

Cass, D., "Optimum Growth in an Aggregative Model of Capital Accumulation," *Review of Economic Studies*, 32, July 1965, 233-240.

De Long, J.B., "Productivity Growth, Convergence, and Welfare: Comment," *American Economic Review*, 78, December 1988, 1138-1154.

Greenwood, J. and B. Jovanovic, "Financial Development, Growth, and the Distribution of Income," unpublished, University of Western Ontario, March 1989.

King, Robert G. and S. Rebelo, "Transitional Dynamics and Economic Growth in Neoclassical Economies," unpublished, University of Rochester, July 1989.

Koopmans, T.C., "On the Concept of Optimal Economic Growth," in *The Econometric Approach to Development Planning*, North Holland, Amsterdam, 1965.

Romer, P.M., "Capital Accumulation and Long-Run Growth," in R.J. Barro, ed., *Modern Business Cycle Theory*, Harvard University Press, Cambridge MA, 1989.

Solow, R.M., "A Contribution to the Theory of Economic Growth," *Quarterly Journal of Economics*, 70, February 1956, 65-94.

Summers, R. and A. Heston, "A New Set of International Comparisons of Real Product and Price Levels: Estimates for 130 Countries," *The Review of Income and Wealth*, 34, March 1988, 1-25.

White, H., "A Heteroskedasticity-Consistent Covariance Matrix Estimator and a Direct Test for Heteroskedasticity," *Econometrica*, 48, May 1980, 817-838.

Table 1

Cross-State Regressions (48 observations)

	Sample	CONST.	$\log(y_0)$	STRUCTURE	R^2	σ
Personal Income:						
1.	1929-88	.035	-.0107	--	.84	.0018
		(.001)	(.0007)			
Gross State Product (GSP):						
2.	1963-86	.057	-.0172	--	.38	.0040
		(.007)	(.0032)			
3.		.053	-.0205	0.69	.48	.0037
		(.007)	(.0032)	(0.24)		
4.		.060	-.0224	--	.46	.0038
		(.007)	(.0037)			
5.		.055	-.0224	0.53	.50	.0037
		(.007)	(.0036)	(0.28)		
6.	1965-70	.095	-.0338	--	.30	.0089
		(.017)	(.0076)			
7.		.062	-.0272	1.05	.49	.0077
		(.017)	(.0068)	(0.26)		
8.	1970-75	.099	-.0356	--	.10	.0161
		(.038)	(.0159)			
9.		.082	-.0358	2.03	.51	.0120
		(.028)	(.0119)	(0.33)		
10.	1975-80	-.017	.0166	--	.03	.0141
		(.035)	(.0142)			
11.		.045	-.0235	1.91	.78	.0069
		(.018)	(.0077)	(0.16)		
12.	1980-85	.015	-.0540	--	.27	.0154
		(.003)	(.0132)			
13.		.084	-.0363	1.91	.65	.0108
		(.026)	(.0095)	(0.27)		

Table 1, continued

	Sample	CONST.	$\log(y_0)$	STRUCTURE	R^2	σ
Pooled 5-year samples:						
14.	1965-85	indiv.	-.0295 (.0066)	--	.16	.0144
15.		indiv.	-.0294 (.0044)	1.84 (0.12)	.62	.0096
16.		indiv.	-.0312 (.0046)	indiv.	.64	.0095

Notes to Table 1: Standard errors of coefficient estimates appear in parentheses. Estimation is by OLS, but estimated standard errors are similar with White's (1980) heteroskedasticity-consistent estimator. The dependent variable is $(1/T)\cdot\log(y_T/y_0)$, where y_T and y_0 are real per capita state gross state product (or real personal income for regression 1) at the end and beginning of the sample, respectively. STRUCTURE is an estimate of the state growth rate if each of the state's sectors had grown at the national average rate. Regressions 14-16 are pooled results for the 5-year sub-samples, 1965-70, 1970-75, 1975-80, 1980-85. Each regression contains individual constants for each sub-sample. Regression 16 allows also for individual coefficients for the variable STRUCTURE.

Table 2

Standard Deviations (Unweighted) of Logarithms of Product and Income (σ_t)

Year	(1) 48 States (GSP)	(2) 48 States (Income)	(3) 114 Countries (GDP)	(4) 60 Countries (GDP)	(5) 20 OECD (GDP)
1929	--	.37	--	--	--
1933	--	.39	--	--	--
1940	--	.36	--	--	--
1945	--	.24	--	--	--
1948	--	.22	--	--	--
1950	--	.24	--	.89	.60
1955	--	.22	--	.91	.55
1960	--	.20	.93	.93	.53
1963	.18	.18	--	--	--
1965	.17	.18	.99	.95	.50
1970	.15	.16	1.02	.98	.45
1975	.14	.13	1.05	.99	.42
1980	.17	.13	1.11	1.03	.42
1985	.15	.15	1.15	1.06	.43
1986	.14	.16	--	--	--
1988	--	.17	--	--	--

Notes to Table 2: Column 1 applies to real per capita gross state product (GSP), column 2 to real per capita state personal income, columns 3-5 to real per capita gross domestic product (Summers and Heston, 1988). The 20 original OECD countries included in the OECD sample are Austria, Belgium, Canada, Denmark, France, Germany, Greece, Iceland, Ireland, Italy, Luxembourg, the Netherlands, Norway, Portugal, Spain, Sweden, Switzerland, Turkey, the United Kingdom, and the United States. These countries were members as of the agreement in December 1960, which came into effect in September 1961. Subsequent members, not included in the sample, are Japan (1964), Finland (1969), Australia (1971), and New Zealand (1973).

Table 3

Cross-Country Regressions (114 countries)

	Sample	CONST.	$\log(GDP_0)$	PRIM	SEC	R^2	σ
1.	1960-85	.019 (.002)	.0047 (.0019)	--	--	.05	.0191
2.		-.010 (.005)	-.0090 (.0028)	.037 (.007)	.023 (.013)	.33	.0161
3.	1960-65	.028 (.003)	.0091 (.0032)	--	--	.07	.0315
4.		.026 (.009)	.0082 (.0056)	.003 (.013)	.001 (.025)	.07	.0318
5.	1965-70	.028 (.003)	.0035 (.0027)	--	--	.02	.0282
6.		-.005 (.007)	-.0099 (.0041)	.048 (.011)	.011 (.020)	.20	.0256
7.	1970-75	.020 (.003)	.0048 (.0031)	--	--	.02	.0332
8.		-.013 (.010)	-.0052 (.0053)	.042 (.013)	.009 (.021)	.11	.0319
9.	1975-80	.016 (.003)	.0080 (.0029)	--	--	.06	.0324
10.		-.004 (.010)	-.0078 (.0052)	.018 (.014)	.036 (.020)	.11	.0318
11.	1980-85	-.004 (.004)	.0048 (.0029)	--	--	.02	.0342
12.		-.048 (.010)	-.0166 (.0050)	.043 (.013)	.070 (.020)	.22	.0309
Pooled 5-year samples:							
13.		indiv.	.0060 (.0013)	--	--	.14	.0319
14.		indiv.	-.0056 (.0023)	.029 (.006)	.028 (.009)	.21	.0307
15.		indiv.	-.0057 (.0023)	indiv.	indiv.	.22	.0307

Notes to Table 3: Standard errors of coefficient estimates appear in parentheses. Estimation is by OLS, but estimated standard errors are similar using White's (1980) estimator. The dependent variable is $(1/T)\cdot\log(GDP_T/GDP_0)$, where GDP_T and GDP_0 are real per capita gross domestic product (from Summers and Heston, 1988) at the end and beginning of the sample, respectively. PRIM is the primary-school enrollment rate and SEC is the secondary-school enrollment rate. The regressions that start in 1960 or 1965 use the 1960 values, and those that start later use the 1970 values. These data come from UNESCO *Yearbooks* and the World Bank (see Barro, 1989b). Regressions 13-15 are pooled results for the 5-year sub-samples, 1960-65, 1965-70, 1970-75, 1975-80, 1980-85. These regressions allow for individual constant terms for each sub-period. Regression 15 allows also for individual coefficients on the variables PRIM and SEC.

13

Table 4

Cross-Country Regressions (OECD sample, 20 countries)

	Sample	Const.	$\log(GDP_0)$	PRIM	SEC	R^2	σ
1.	1960-85	.041	-.0085	--	--	.45	.0051
		(.003)	(.0022)				
2.		.023	-.0124	.016	.012	.60	.0046
		(.008)	(.0028)	(.007)	(.007)		
3.	1960-65	.055	-.0110	--	--	.18	.0127
		(.008)	(.0055)				
4.		.023	-.0189	.027	.026	.30	.0124
		(.023)	(.0074)	(.020)	(.020)		
5.	1965-70	.072	-.0226	--	--	.46	.0126
		(.009)	(.0058)				
6.		.022	-.0281	.051	.005	.65	.0108
		(.020)	(.0071)	(.018)	(.018)		
7.	1970-75	.058	-.0157	--	--	.38	.0092
		(.008)	(.0047)				
8.		.043	-.0173	.011	.010	.41	.0096
		(.027)	(.0057)	(.022)	(.015)		
9.	1975-80	.027	.0009	--	--	.00	.0116
		(.012)	(.0064)				
10.		-.015	-.0052	.028	.033	.23	.0108
		(.031)	(.0069)	(.025)	(.017)		
11.	1980-85	.003	.0043	--	--	.05	.0085
		(.010)	(.0046)				
12.		.001	.0031	.000	.006	.06	.0090
		(.026)	(.0059)	(.021)	(.015)		
Pooled 5-year samples:							
13.		indiv.	-.0101	--	--	.47	.0118
			(.0026)				
14.		indiv.	-.0137	.022	.014	.51	.0115
			(.0031)	(.010)	(.008)		
15.		indiv.	-.0136	indiv.	indiv.	.58	.0111
			(.0031)				

Note: See the notes to Table 3.

Table 5

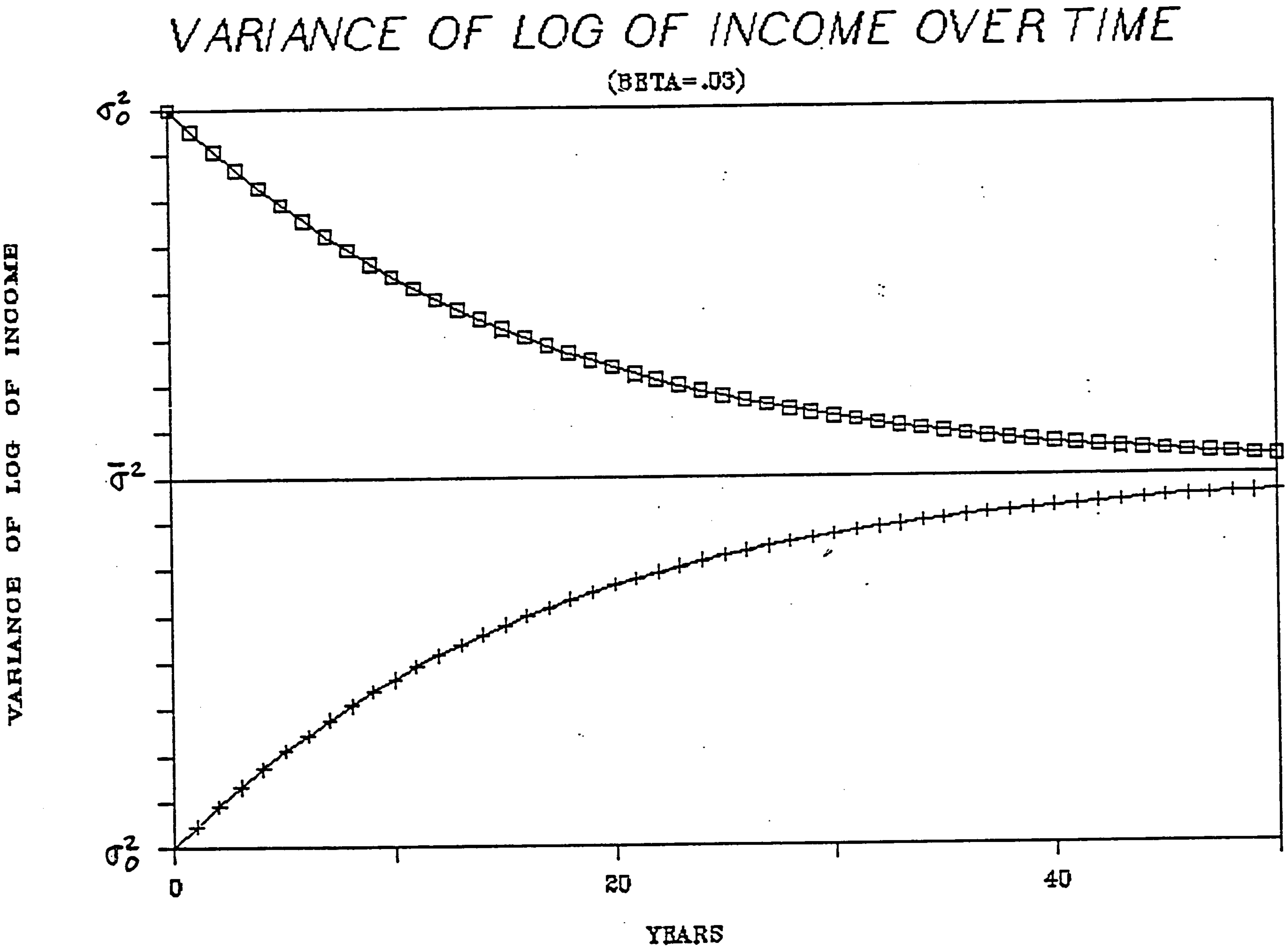
Cross-State Regressions with Additional Variables

	(1)	(2)	(3)	(4)
Constant	.060 (.007)	.055 (.007)	.052 (.008)	.062 (.008)
$\log(GSP_{1963})$	-.0224 (.0037)	-.0224 (.0036)	-.0218 (.0037)	-.0228 (.0035)
COLLEGE	.107 (.044)	.056 (.051)	.045 (.052)	.050 (.049)
STRUCTURE	--	.53 (.28)	.59 (.29)	.76 (.28)
PUBI	--	--	.087 (.086)	.140 (.083)
GOV	--	--	--	-.080 (.031)
R^2	.46	.50	.51	.58
σ	.0038	.0037	.0037	.0035

Note: The dependent variable is the growth rate of real per capita gross state product from 1963 to 1986. COLLEGE is the fraction of the adult population in 1960 that had completed 4-year college. PUBI is the average from 1950 to 1962 of the ratio of state government capital expenditures to state personal income. GOV is the average from 1963 to 1986 of the ratio of expenditures by state and local governments (exclusive of transfers from state to local governments) to gross state product. See the notes to Table 1 for additional information.

30

Figure 1


VARIANCE OF LOG OF INCOME OVER TIME
(BETA=.03)
VARIANCE OF LOG OF INCOME
σ_0^2
$\bar{\sigma}^2$
σ_0^2
0
20
40
YEARS

31

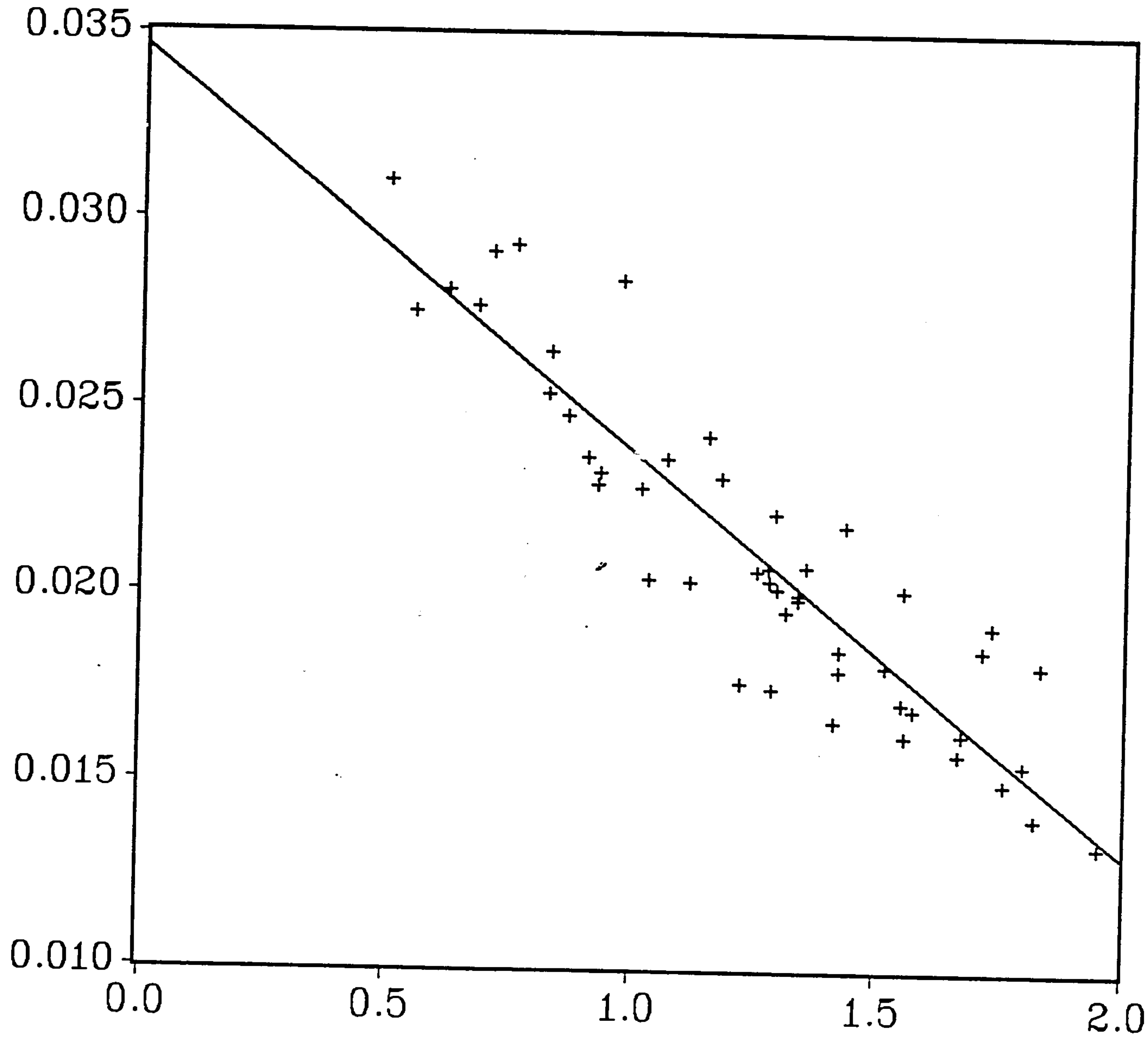


Figure 2 Growth Rate of per capita Real Personal Income (1929–88) versus $\log(y_{1929})$ for 48 States

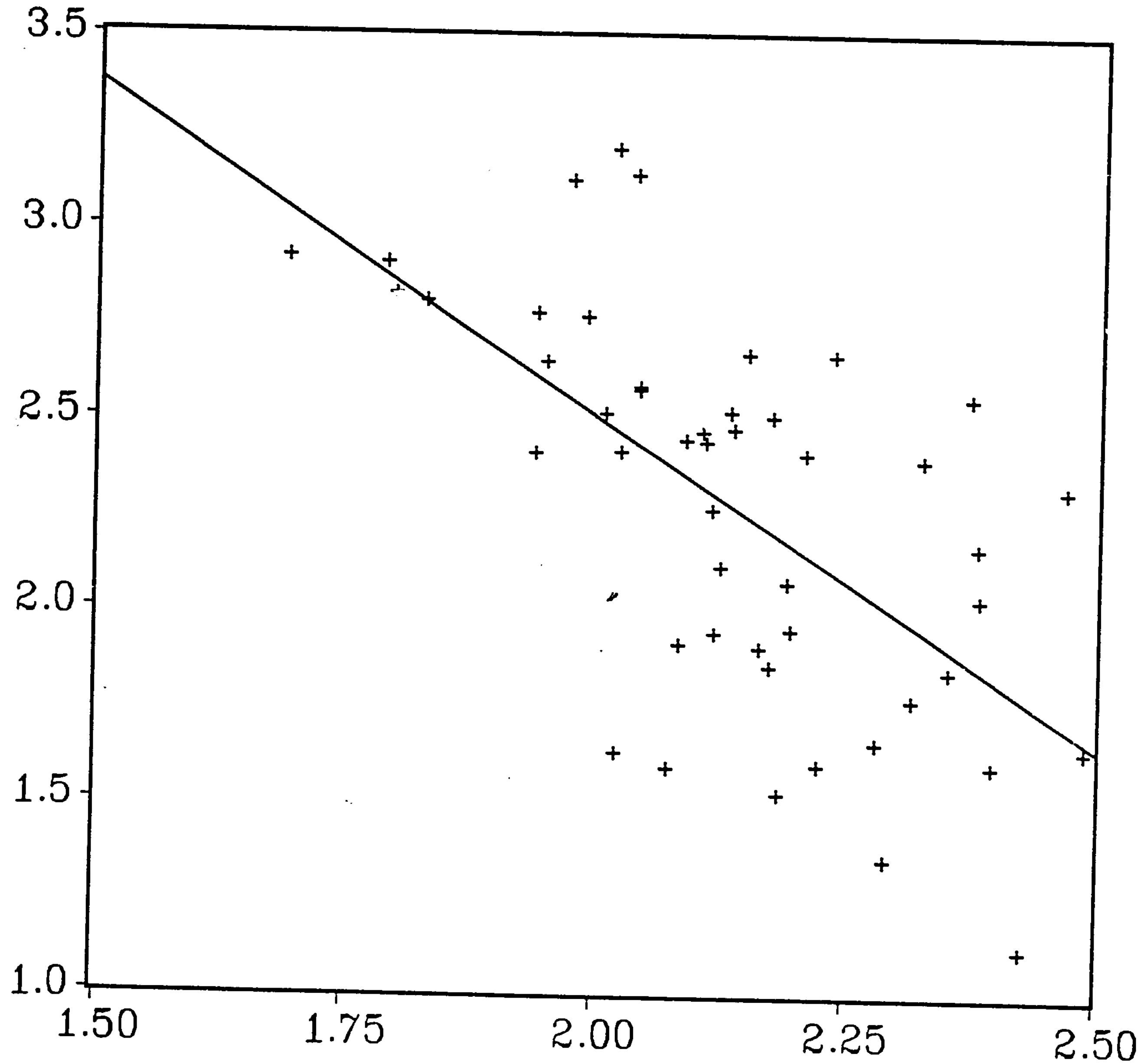


Figure 3 Growth Rate of per capita GSP (1963-86) versus log(GSP_{1963}) for 48 States

33

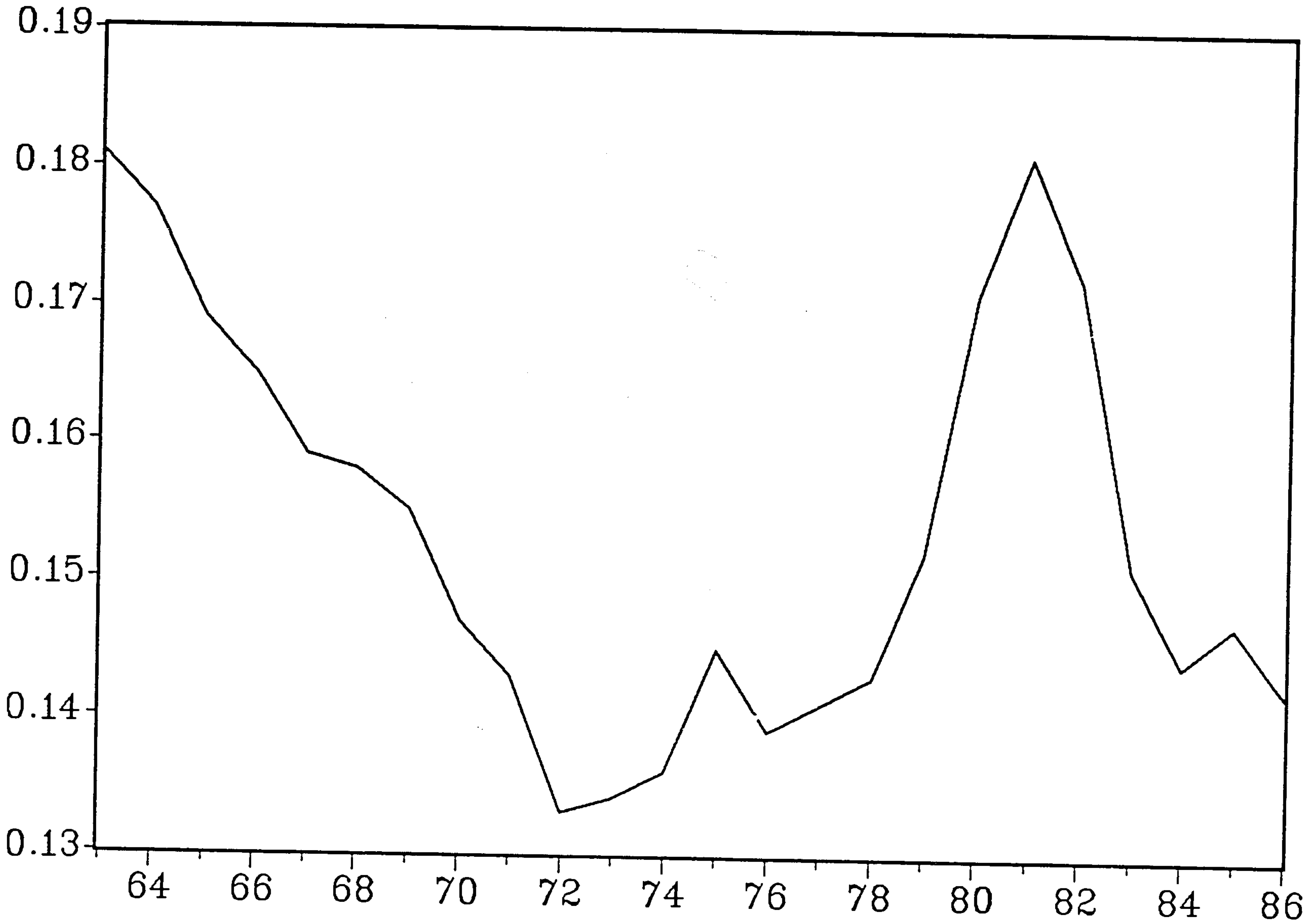


Figure 4 Standard Deviation of log(GSP) across 48 States

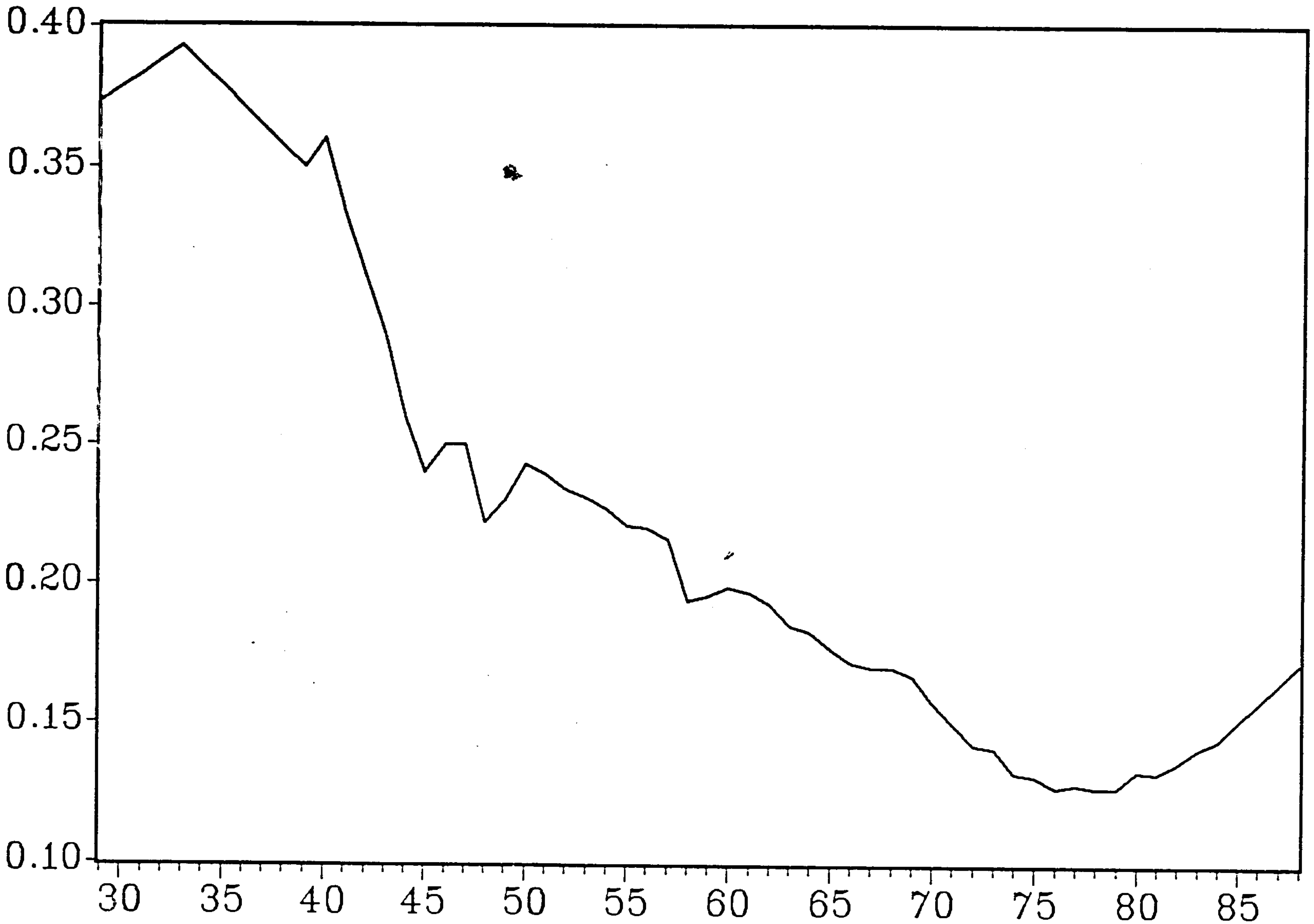


Figure 5 Standard Deviation of log(real per capita personal income) across 48 States (Data Interpolated for 1930–32, 1934–38, 1941–42)

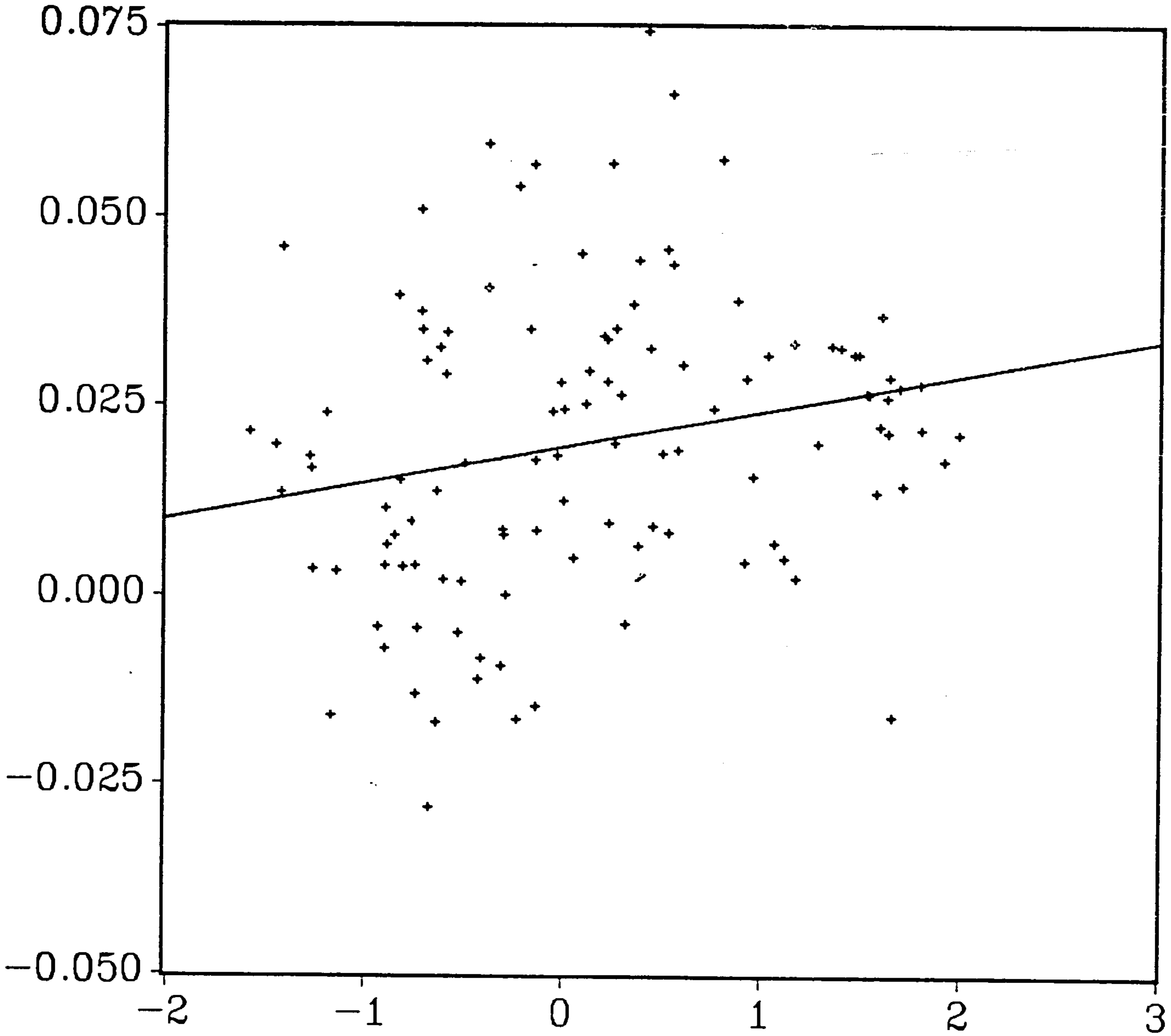


Figure 6 Growth Rate of per capita GDP (1960−85) versus $\log(GDP_{1960})$ for 114 Countries

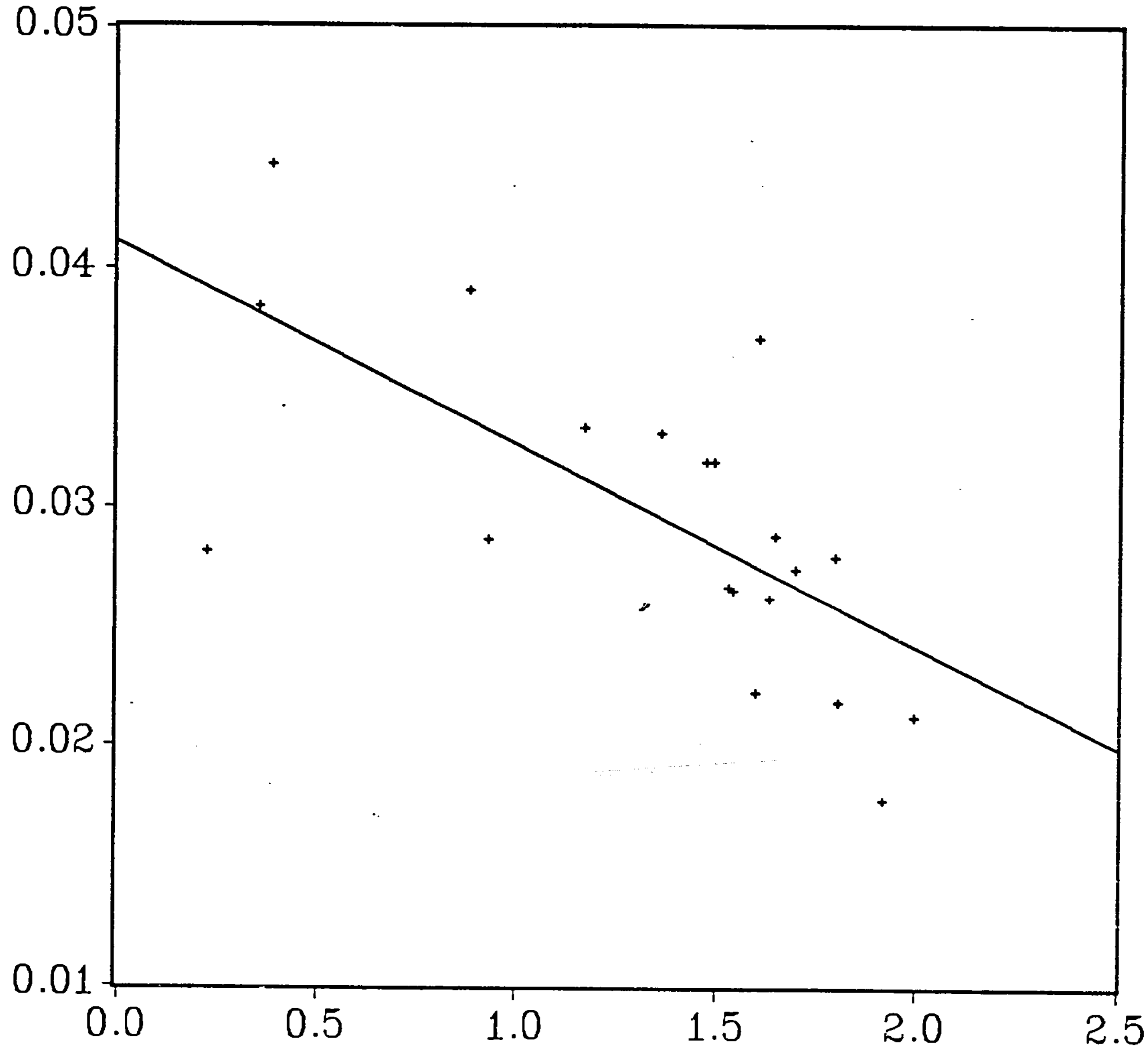


Figure 7 Growth Rate of per capita GDP (1960–85) versus $\log(GDP_{1960})$ for 20 OECD Countries