Income convergence among U.S. states: cross-sectional and time series evidence

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Abstract. We perform convergence tests on the U.S. states for per capita income from 1930 to 2009. Cross-sectional tests support overall σ-convergence and β-convergence but may not hold true for the last three decades. Time series tests suggest that about half of the states exhibit stochastic convergence and of these all are also β-converging. Probit regressions reveal that the likelihood a state is converging is a function of changing capital to labour ratios, the size of the agricultural sector, and levels of taxation and tax revenue. Regional disparities in convergence remain among the southern and midwestern states. JEL classification: R1

Convergence des niveaux de revenus entre les états américains : résultats á partir de donnés transversales et de séries chronologiques. On présente des tests de convergence des revenus per capita des états américains entre 1930 et 2009. Les tests transversaux supportent l'hypothèse de convergence de type σ et de type β, mais pourraient ne pas tenir pour les trois dernières décennies. Les tests utilisant les séries chronologiques suggèrent que la moitié des états montrent une convergence stochastique, et que de ceux-ci tous supportent l'hypothèse de convergence de type β. Des régressions par la méthode probit révèlent que la vraisemblance qu'un état converge dépend des rapports capital/travail changeants, de la taille du secteur agricole, et des niveaux de taxation et de rentrées fiscales. Des disparités régionales dans la convergence demeurent entre les états du sud et du midwest.

1. Introduction

One object of investigation among growth economists is convergence of various economies to a common steady-state level of income. Neoclassical growth models, assuming declining marginal productivity of capital, predict convergence (conditional on labour growth and savings rate), whereas endogenous growth

Allin Cottrell contributed to an earlier version of this paper. I thank him for his programming assistance in gretl. Email: heckeljc@wfu.edu

Canadian Journal of Economics / Revue canadienne d'Economique, Vol. 46, No. 3
August / août 2013. Printed in Canada / Imprimé au Canada

0008-4085 / 13 / 1085–1109 /

models often predict divergence. One problem in testing these predictions is how exactly to measure convergence. The empirical literature contains at least three conceptions of convergence. The notion of σ-convergence requires that the dispersion of incomes across economies falls over time. A related but distinct concept is that of β-convergence, which requires that economies with lower initial incomes grow at a faster rate than economies with higher initial incomes. Finally, stochastic convergence requires that shocks to relative income (i.e., the economy's income relative to the average of all economies) must have only temporary effects.

Cross-national tests for convergence typically examine the case for conditional, rather than absolute, convergence. Only once differences in institutional structure, policy variables, demographics, and other characteristics are controlled, is evidence for convergence discovered. This weakened form of convergence means not that nations are literally converging in income, but only that they could be if they all had a similar set of institutions and demographic characteristics. Absolute convergence would be expected to occur only among a group of economies that operate under similar conditions. Arguably, the individual states of the U.S. form such a group: they are all part of the same federal system and subject to a relatively consistent set of economic and political institutions. There are few barriers to mobility of labour and capital in search of the highest returns. And although all states are not identical in terms of regulations and demographic characteristics, these differences are not as great as they are for various groups of nations.

Several studies have tested for cross-sectional β-convergence of the U.S. states (or 'catch-up') by regressing growth rates against initial income levels. Most find evidence in favour of 'catch-up', but by including additional explanatory variables, they are actually supporting only conditional convergence. In constrast, Barro and Sala-i-Martin (1991) show absolute β-convergence holds over the span 1880–1988 and within each decade as well except for the 1920s and 1980s. σ-convergence for the states has received much less attention.

Because the cross-sectional approach does not allow for distinctions among the states, time-series tests for stochastic and β-convergence have been advocated by Carlino and Mills (1993), but their study and many others tend to focus only on convergence among the regions.[1] Examining individual states for convergence offers two potential improvements over analysis of regional convergence. First, even if regional convergence occurs, there can still be significant divergence within a region (Carlino and Mills 1996). Second, with only a handful of U.S. regions it is not practical to run tests for the *determinants* of convergence or divergence. Scholars are left to speculate on the unique features of the few regions that are found to differ from the others in their stochastic properties. Because there are far more states than regions, there is likely to be more variation among the convergers

1 Carlino and Mills (1996) test for convergence of states to their respective regional averages, but if the regions have not fully converged (as indicated by the evidence presented in Carlino and Mills (1993), Tomljanovich and Vogelsang (2002), and others), this is still not the same as full convergence across the nation.

and divergers, thus allowing for regression analysis of the determinants of state convergence, as in the present study. This may assist policymakers in their efforts to further the convergence process; from a policy perspective, simply knowing whether or not convergence is occurring is not sufficient.

In this paper we test for σ-, β-, and stochastic convergence among the U.S. states over the period 1930–2009, as well as the determinants of convergence. The rest of the paper proceeds as follows. Section 2 presents cross-sectional tests for state convergence, which supports σ-convergence and β-convergence over the 80-year period as a whole, but not for the last three decades. Time series tests in section 3 suggest that 23 of the 49 continental states (including District of Columbia) exhibit stochastic convergence and of these all are in the process of, or have already achieved, β-convergence. Section 4 reveals that, in particular, states with larger reductions (or smaller increases) to their capital to labour ratios, smaller agricultural sectors, and less government intervention in terms of taxation and tax revenue are more likely to exhibit evidence of convergence. In addition, midwestern states are significantly more likely to be converging, especially compared with southern states. The final section concludes the paper.

2. Cross-sectional tests for convergence

As mentioned above, σ-convergence denotes a process resulting in a reduction in the dispersion of income over time, whereas β-convergence denotes a process whereby economies that start below (respectively, above) the average income level tend to grow at a faster (respectively, slower) than average rate. The presence or absence of one type of convergence does not yield any information regarding the other (Quah 1996). For example, the former can occur without the latter under certain types of club convergence, when only subsets of economies converge among themselves. The reverse can hold true if – to give an extreme example – relative income levels become exactly reversed from their starting positions. Furthermore, as noted by Quah (1993a, 1993b, 1996), cross-sectional regressions represent average behaviour of the economies in question and do not shed direct light on the behaviour of the entire distribution, which he argues is more important. We first show in this section that the cross-sectional test results for σ-convergence and β-convergence for the U.S. states are consistent with each other before we turn in the next section to consider each state in isolation.

To test for each type of convergence, the log of per capita income for each state and District of Columbia (DC), taken from the Bureau of Economic Analysis (BEA) website, is analyzed for the years 1930–2009.[2] For simplicity, the units of observation will be referred to as states hereinafter despite the inclusion of

2 Ideally, convergence would be tested using real state values that incorporate price differences across the states, such as in Izraeli and Murphy (1997). Nominal values are used here because a state price index series does not go back far enough. Other studies that rely on nominal income include Barro and Sali-i-Martin (1991, 1992), Crown and Wheat (1995), Carlino and Mills

DC. The BEA measure divides the total state income for the year by the midyear population estimate. Per capita income series for Alaska and Hawaii do not begin until 1950 and are not included except where noted.

2.1. σ-convergence

Common measures of dispersion include the standard deviation and coefficient of variation (CV), which divides the standard deviation by its mean. The latter is more appropriate for a time comparison when the mean value has drift.[3] Kenworthy (1999) showed that the coefficient of variation (CV) for per capita state income did not reveal a distinct pattern across four separate years spread out over 1969–1996, but an annual series was not presented. Crown and Wheat (1995) do present a table of annual CVs covering 1929–1993, but remark on the trends only in passing. They conclude in favour of long-run convergence based solely on visual inspection of the table. Yet the underlying data are from mixed sources in different years, making valid comparisons difficult. Furthermore, neither study includes any significance testing for the CV data.

Figure 1 graphically displays the annual CV across the states from 1930 to 2009 using the log form of per capita personal income. The trend is clearly for dispersion to fall over time and the time trend is statistically significant, as shown in the following regression equation (Newey-West heteroscedasticity-autocorrelation robust t-statistics in parentheses).

$$\widehat{CV}_t = \underset{(14.34)}{0.091} \; \underset{(-11.05)}{-0.019 \log(t)} \qquad R^2 = .91.$$

The log form for time trend shown in the graph and regression equation fits better than a linear trend, indicating that annual reductions in dispersion are diminishing over time. None of these results is significantly influenced by the addition of Alaska and Hawaii starting in 1950. The conclusion is that σ-convergence occurs for state per capita income, suggesting that the distribution of income across the states is becoming more uniform, consistent with Crown and Wheat's (1995) observation based on mixed data sources through 1993.

Although there is an overall declining dispersion of income, the trend is not consistent. Over an almost 50-year period from the mid-1930s until 1980, dispersion in state per capita incomes continued to fall.[4] Thereafter, the degree of dispersion reversed trend and began to increase for several years, and the slight reductions in dispersion that began after the end of the Reagan years were not great enough to return to the previous trend. Because Crown and Wheat ended

(1996), and Johnson (2000). Others, such as Evans (1997), deflate all state values by the annual GNP deflator. Our results remain the same when we convert to real values using the GNP deflator.

3 Although we present results for the CV, the graph and regression reported below are very similar if the standard deviation is used instead.

4 Crown and Wheat's data suggest that a very slight increase also occurs in 1979. In both data sets, 1981 shows a lower CV than in 1980, with annual increases from 1982–1988.

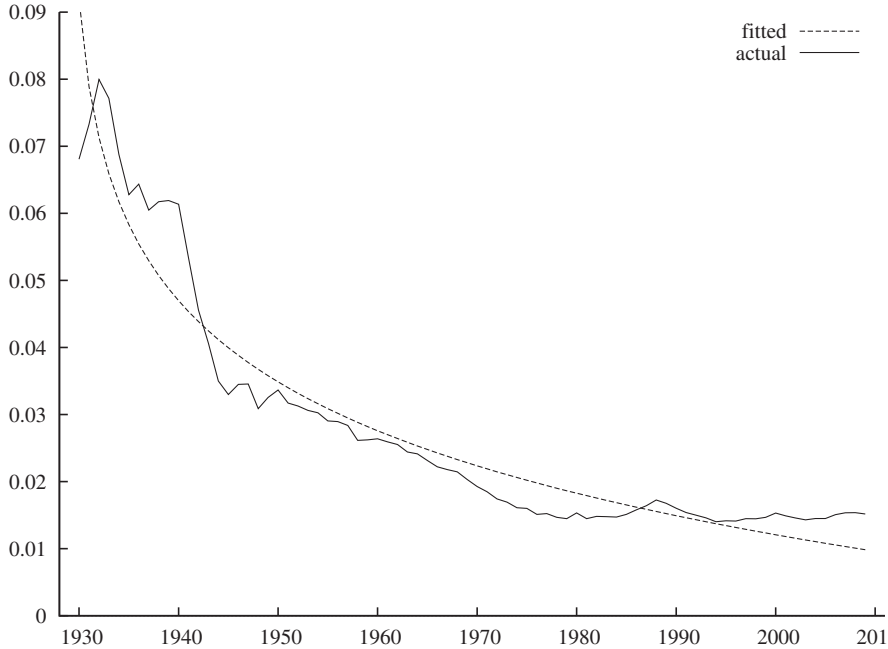

FIGURE 1 σ -convergence for state per capita income, 1930–2009

their sample in 1993, with the last five years showing annual declines in CV, the impression might be that the lack of continued σ-convergence in the 1980s was an anomaly. Extending the data here shows that the slight decline in annual CVs starting again in 1989 had ended by 1996, just three years after Crown and Wheat's sample. That σ-convergence was not continuing during the most recent three decades is confirmed by lack of significance of a time trend (either log or linear) when the sample is limited to the 1980–2009 period (p-value = 0.33 or 0.27, respectively). Reducing the sample further to only the post-1988 years does not improve the time trend p-values. Thus, we conclude that income dispersion fell significantly over time, although the process of σ-convergence appears to have ended in 1980.

2.2. β-convergence

Barro and Sala-i-Martin (1991) derived a formula, based on a Solow-growth model, to capture the rate of economic convergence over time. Their concern, different from the notion of σ-convergence described in the previous section, was to determine how fast poorer states were catching up to wealthier states. Their empirical specification took the form

$$(1/T)(y_{it} - y_{i,t-T}) = \alpha - (1 - e^{-\beta T})(1/T)y_{i,t-T} + \varepsilon_{it}, \tag{1}$$

TABLE 1
Cross-sectional tests for β-convergence on per capita income

	Eq. (1) $0 < \beta < 1$	Eq. (2) $\tilde{\beta} < 0$
1930–2009	0.017*	– 0.009*
	(9.05)	(– 18.82)
1930–1940	0.012*	– 0.012*
	(4.13)	(– 4.39)
1940–1950	0.053*	– 0.041*
	(9.09)	(– 11.96)
1950–1960	0.023*	– 0.021*
	(6.36)	(– 7.16)
1960–1970	0.026*	– 0.023*
	(5.57)	(– 6.35)
1970–1980	0.014*	– 0.013*
	(2.85)	(– 3.06)
1980–1990	0.011	– 0.010
	(1.33)	(– 1.40)
1990–2000	0.008	– 0.008
	(1.76)	(– 1.83)
2000–2009	0.012	– 0.012
	(1.39)	(– 1.47)
1930–2009	0.024*	– 0.021*
(pooled)	(12.71)	(– 14.37)

NOTES: t-statistics in parentheses. All regressions include a constant term. Equation (1) estimates produced by Iterative Non-Linear Least Squares. Equation (2) estimates produced by Ordinary Least Squares. 'Pooled' indicates results from pooled per decade data with cross-equation restrictions. * = significant at 5% level.

where y_{it} is log per capita income of state i in year t, covering T total years, ε_{it} are exogenous shocks, and β represents the annual rate of convergence. Convergence requires $0 < \beta < 1$. They found state economies for the continental U.S. were converging over the period 1880–1988 at the rate of roughly 1.7% per year.

Following from Barro and Sala-i-Martin (1991), the first column of table 1 presents non-linear least squares estimates of β for the full time period and by decade. The full sample period supports the finding of β-convergence at about the same 1.7% annual rate, over the 80-year period studied here. Barro and Sala-i-Martin also find that convergence rates vary by decade but, except for the 1920s and 1980s, always generate estimates consistent with β-convergence. The results here (which also include District of Columbia, and Alaska and Hawaii beginning in 1950) suggest that the process of convergence ended at the end of the 1970s; although each decade generates a β value within the relevant range, the decades of the 1980s, 1990s, and 2000s do not generate statistically significant coefficients. Barro and Sala-i-Martin estimate a negative (but not statistically significant) coefficient estimate for the 1980s – on a sample that ends in 1988,

and does not include Alaska, Hawaii, or District of Columbia. Dropping these three states from our sample does not affect any of our reported results.

Finally, also following from Barro and Sala-i-Martin, in the bottom row of table 1 the separate per decade equations are pooled in a joined system of equations with cross-equation restrictions for a single estimated value of β (the intercept is allowed to vary across decades). An iterative non-linear least squares routine is utilized to estimate a common convergence rate of 2.4% per annum. Again, restricting the sample to only the 48 continental states does not alter the estimates.

Alternatively, a simpler form of equation (1) often employed by others (e.g., Izraeli and Murphy 1997; Miller and Russek 1997) is

$$1/T(y_{it} - y_{i,t-T}) = \alpha + \tilde{\beta} y_{i,t-T} + \varepsilon_{it}, \tag{2}$$

where $\tilde{\beta} = -(1 - e^{-\beta T})(1/T)$. In this case, convergence requires $\tilde{\beta} < 0$ and can be estimated by Ordinary Least Squares. Coefficient estimates reported in the second column of table 1 are always negative, but again not statistically significant for the 1980s, 1990s, or 2000s, largely mirroring the findings from equation (1).

The cross-sectional regressions assume that exogenous shocks are random and independently distributed. But if shocks are regionally related, the disturbance terms would be spatially correlated. Drawing on the spatial growth literature,[5] we test for a variety of potential spatial effects.

Let y_t represent the vector of log state income levels at time t, and let ε_t represent the vector of state errors. The spatial lag model assumes that state growth is affected by growth in nearby states and can be represented by

$$1/T(y_t - y_{t-T}) = \alpha + \tilde{\beta} y_{t-T} + \theta \mathbf{W}(y_t - y_{t-T})/T + \varepsilon_t, \tag{3}$$

where $\mathbf{W}$ is a 49×49 matrix[6] of spatial weights (with zeros on the main diagonal) and θ is a parameter that can be used to test for the importance of the spatial lag.

Alternatively, the spatial cross-regressive model assumes that state growth is affected by the initial incomes of nearby states. The specification

$$1/T(y_t - y_{t-T}) = \alpha + \tilde{\beta} y_{t-T} + \psi \mathbf{W} y_{t-T} + \varepsilon_t \tag{4}$$

is used to represent this case. A significant estimate for ψ indicates the presence of spatial effects.

5 See Fingleton and López-Bazo (2006) for a recent survey.

6 Because the spatial regressions use data beginning in 1930, Alaska and Hawaii are not included in the sample.

TABLE 2
β-convergence on per capita income with spatial effects, 1930–2009

	Weighting matrix					
	State border			City distance		
$\tilde{\beta}$ (convergence term)	− 0.010*	− 0.010*	− 0.009*	− 0.009*	− 0.009*	− 0.009*
	(− 17.91)	(− 17.86)	(− 18.45)	(− 18.62)	(− 18.62)	(− 18.43)
θ (spatial lag)	− 0.003			0.000		
	(− 1.33)			(0.06)		
ψ (spatial cross-regressive)		− 0.000			0.000	
		(− 1.24)			(0.02)	
λ (spatial error)			0.033			0.091
			(0.49)			(0.57)

NOTES: t-statistics in parentheses. * = significant at 5% level.

Finally, if shocks to one state affect growth rates in other states, then $\varepsilon_t = \lambda \mathbf{W} \varepsilon_t + \upsilon_t$, in which case

$$\begin{aligned} 1/T(y_t - y_{t-T}) &= \alpha + \tilde{\beta} y_{t-T} + \lambda \mathbf{W} \varepsilon_t + \upsilon_t \\ &= \alpha + \tilde{\beta} y_{t-T} + \lambda \mathbf{W}(y_t - y_{t-T})/T - \lambda \mathbf{W} \alpha - \lambda \mathbf{W} y_{t-T} + \upsilon_t, \end{aligned} \tag{5}$$

which can be estimated via non-linear least squares with the appropriate restrictions on λ.

Typically, the weighting matrix reflects the notion that conditions within any given state are expected to have a larger impact on nearby states than on distant ones. Following Coughlin, Garrett, and Hernández-Murillo (2004), we consider two potential weighting matrices. In the first case $\mathbf{W}$ is composed of dummy variables that are equal to 1 for states that share a border, 0 otherwise.[7] The alternative weighting matrix has W_{ij} proportional to the reciprocal of the distance between the most populous cities in states i and j. Estimates for equations (3)–(5) are presented in table 2, covering the period 1930–2009. In each case we are unable to reject the null of no spatial effects. This applies for both weighting matrices. This is confirmed using the traditional Moran's I test for spatial dependence, which does not distinguish among the types of potential spatial correlation, and LM tests specific for spatial lag or spatial cross-regressive effects, as shown in table 3.[8] In addition, the inclusion of spatial terms does not affect the estimated values for $\tilde{\beta}$, further indicating that our previous estimates are likely free from spatial bias.

Thus, we conclude that β-convergence is observed over the period 1930–2009 (as well as for the shorter period 1950–2009, for which the data set includes Alaska and Hawaii). However, these findings are dominated by β-convergence only through the 1970s.

7 Rey and Montourri (1999) also use the border dummy approach for state spatial effects.
8 Details on computing the test statistics are presented in Greene (2008, 218–19).

TABLE 3
Spatial correlation test statistics

	Weighting matrix	
	State border	City distance
Moran's I for spatial correlation	0.575	0.160
LM test for spatial lag	0.000	0.000
LM test for spatial error	0.206	0.030

NOTES: Null is no spatial effect. Moran's I is the transformed z-statistic. LM statistics are distributed as chi-square with one degree of freedom.

2.3. Summarizing the cross-sectional results

The cross-sectional results suggest convergence is found overall for the full sample period using either notion of convergence: σ-convergence or β-convergence. One potential concern, as noted above, is that the traditional cross-sectional approach treats all states the same. In essence, the implicit null is that all states are converging, which is tested against the implicit alternative of no states converging. Strict reliance on these measures could give a misleading impression of the true distribution dynamics if some states are converging when others are not. For this reason, Quah (1996) presented kernel estimates from a Markov transition matrix to consider the transitional dynamics of the entire distribution. He was particularly concerned with the potential emergence of endogenous convergence clubs based on initial income conditions. His international sample for 1962–1984 did not show much transition out of the original income groups. In particular, the ergodic distribution became bimodal, suggesting two convergence clubs where the rich nations become richer, the poor nations become relatively poorer, and the the middle class shrinks. He interpreted these results as evidence of convergence within clubs but divergence (or 'polarization') across the two clubs. In contrast, his sample of U.S. states from 1948 to 1989 showed strong evidence of convergence with much greater mobility across relative income groups. He further noted that the ergodic distribution appears hump-shaped, suggesting the absence of convergence clubs among the states. That all the tests (σ-convergence, β-convergence, and Quah's Markov transition) are consistent with convergence among the states over the entire respective sample periods suggests that distributional concerns among the states may be less important for determining convergence than for an international sample. Yet results presented here suggest that the finding of σ-convergence and β-convergence did not continue beyond 1980.[9] Particular sample time periods may be a more important consideration for the U.S. states than which test method or notion of convergense is utilized.

9 Quah did not consider period subsamples, so it is possible he overlooked changes in the timing of the process of convergence if transitions were happening less frequently in the 1980s than earlier.

There are several potential explanations for the seeming end of convergence as indicated by the σ-convergence and β-convergence tests. From a political economy perspective, Olson (1982, 1983) argued that southern and western states did not have as many distributional coalitions[10] influencing policy and institutions that tend to obstruct growth prospects. Thus, these regions were able to grow faster than the midwest and eastern states burdened by such sclerotic groups. But he also posited that over time such groups would continue to form, eventually transfering the south from a 'leader' to a 'laggard' in growth. As such, the south's rapid catching-up would soon be reversed. The particular timing of the end of convergence found here also roughly coincides with the period of the 'Great Moderation,' in which output and employment fluctuations lessened, dampening business cycle amplitudes. Owyang, Piger, and Wall (2008) find that volatility reductions tended to be larger in those states that had previously experienced greater volatility. Stabalization of output would conceivably reduce the range over which states grew. Lower variation in growth rates across states would limit the opportunity for poorer states to catch up.[11] Relatedly, the 1980s fostered a new era of decentralization (Gaffakin and Warf 1993), which has perhaps been reversed again in just the past few years (Conlan and Posner 2012). Less central direction of resources (which is not the same as less total spending) by the federal government gave the states greater freedom to pursue their own policies, which could result in more divergent outcomes. Finally, Quah's notion of endogenous income clubs may have emerged specifically during this period for the above or alternative reasons. We note here only that the conclusion of convergence depends on the period under study, and we leave for future work testing various theories as to why convergence appears to have ended when it did.

The tests discussed so far can be useful to describe general patterns of convergence. However, even if convergence is generally taking place from 1930 to 1980, as confirmed by each of the measures, it can still be true that not every state is part of the convergence process. Even though the states overall have been converging since 1930, some states may not be converging. And even though the states *as a whole* have generally not been converging since 1980, many states might still be converging. For either period, stating that convergence by the states does or

10 Western states were thought to have a late start on the formation of interest groups because they were settled, and granted statehood, at later dates. The southern states had their governing coalitions and institutional structures upset by the Civil War and Reconstructionism.

11 A reviewer offered an alternative explanation for the impact of the Great Moderation on convergence. Focusing on price level stability, the reviewer suggests price differences across states would be more pronounced prior to the 1980s. Thus, there would be reduced differences between nominal and real incomes (in the PPP-adjusted sense) after the onset of the Great Moderation. Lack of convergence in nominal incomes during the Great Moderation may properly indicate a lack of convergence in real incomes. But because nominal income convergence in the earlier periods captures important price differentials, it does not properly reflect what could be a lack of real income convergence. Thus, although nominal incomes were converging and then stopped converging, the reviewer suggests it could be that real incomes were never converging. Unfortunately, state price indexes do not exist prior to the 1980s, so we are unable to test the reviewer's concern.

does not hold would indicate that all states are, or that no states are, converging. Either conclusion could be potentially grossly misleading. In the next section, we investigate each state individually to determine which particular states are converging or diverging. To do so, we use alternative techniques based on the states' time-series properties.

3. Time-series tests for convergence

The cross-sectional evidence is consistent with overall convergence when the sample is treated as one continuous period but does not establish whether this phenomenon is inclusive of every state. Despite overall convergence, there is still the possibility that certain states were not converging. To determine which particular states are converging or diverging, we now turn to time-series analysis of the individual states. As established by Carlino and Mills (1993), convergence requires two conditions. First, stochastic convergence requires that shocks to relative income (i.e., the state per capita income relative to the average per capita income for all states in the entire nation) must have only temporary effects. Second, as explained above, β-convergence requires that states with initial income below the average grow faster than states with initial income above the average. Both conditions must be met in order to conclude that the state exhibits convergence.

Define log-relative income for state i as

$$\dot{y}_{it} = y_{it} - \bar{y}_t,$$

where $\bar{y}_t$ is the average value of state log per capita income in year t. Because convergence will be tested for each state individually, for ease of notation the subscript i will be dropped from the rest of this section. Note that on this definition a value of zero for log-relative income means that a state is at the national average income per capita.

3.1. Stochastic convergence

The test for stochastic convergence can be implemented by a test of the null hypothesis that log-relative income follows a unit-root process, versus the alternative that it is stationary. If log-relative income does have a unit root, then if a shock pushes the state away from the national average, there is no tendency for it to return, in contradiction to the idea of convergence. This condition is necessary for a state to be considered 'truly' convergent, but that is by no means sufficient for, or even indicative of, other notions of convergence. For example, a state whose relative income fluctuates around a stable mean substantially above or below the national mean satisfies the stationarity criterion yet would fail the β-convergence criterion.

When testing for stationarity, it is necessary to consider a possible trend in log-relative income: for example, a state that is converging toward the average from a starting point substantially below average should show stationary fluctuations around a rising trend. Furthermore, we want to allow (as in, e.g., Tomljanovich and Vogelsang 2002) for the possibility of a break in the convergence trend, that is, the possibility that convergence does not necessarily proceed at a constant average rate. Or in other words, a state's log-relative income may be stationary around a 'broken trend'. This seems reasonable on general grounds, given the historical span of the data. A state might converge rapidly from a starting point far from the average, with convergence slowing as the average is approached.[12]

We therefore employ the methodology of Perron (1997) on 'breaking trend functions' to perform the test for stationarity. Perron's procedure gives the researcher a number of options from which to choose. We used Perron's 'Model 2,' which allows for a shift in both intercept and trend at the time of the break in trend, T_b. Adapting his notation, this model takes the form

$$\dot{y}_t = \mu + \theta \mathrm{D}U_t + \beta t + \gamma \mathrm{D}T_t + \delta \mathrm{D}(T_b)_t + \alpha \dot{y}_{t-1} + \sum_{j=1}^{k} c_i \Delta \dot{y}_{t-j} + e_t, \tag{6}$$

where $\mathrm{D}U_t = I(t > T_b)$, $\mathrm{D}T_t = I(t > T_b)t$, and $\mathrm{D}(T_b)_t = I(t = T_b + 1)$, with $I(\cdot)$ the indicator function. We selected Perron's first method for choosing the break-point, namely, T_b is chosen so as to minimize t_α, the t-statistic for testing the unit root null, $\alpha = 1$.[13] We then selected the lag truncation parameter, k, by utilizing a procedure that 'selects that value of k, say k^*, such that the coefficient on the last lag in an autoregression of order k^* is significant and that the last coefficient in an autoregression of order greater than k^* is insignificant, up to a maximum order k max' (Perron 1997, 359). We set a k max of 8.

For this test we used annual BEA state data for 1930 to 2009. We excluded Alaska and Hawaii, for which comparable data are not available prior to 1950, leaving 48 states plus DC. The sample size for each test is $T = 80$. With these choices, the 5% critical value for t_α is -5.59 (based on Perron's simulations as reported in table 1 of his 1997 paper and using the closest sample size, $T = 70$).

As shown in the first panel of table 4, the unit-root hypothesis was rejected at the 5% significance level for 15 of the 49 states tested. We note that for half of the states the break in trend is estimated to occur between 1939 and 1944 (most of these in 1940 or 1941), although there are nine states for which the break appears to come after 1980. Our results contrast with Choi's (2004) unit root tests, where

12 See also Maddala and Wu (2000) for evidence on changing rates of convergence across a sample of nations.

13 As with the regular Dickey-Fuller test, this is a one-sided test with the critical region in the left tail. Minimizing the t-value therefore means selecting the break-point that is least favourable to the null. When the break-point is selected in this way, obviously the test statistic does not follow the regular t distribution; special critical values must be used, derived by Perron using the Monte Carlo method, and presented in the text below.

TABLE 4
Stochastic convergence tests

	Perron endogeneous break			ADF no break	
State	*t*-statistic	Break year	included lags	*t*-statistic	included lags
AL	– 5.089	1940	0	– 3.718*	8
AR	– 5.040	1950	0	– 2.067	8
AZ	– 4.398	1988	3	– 2.284	2
CA	– 5.225	1943	8	– 2.898	1
CO	– 7.382*	1961	7		
CT	– 4.639	1941	1	– 1.981	1
DC	– 5.443	1940	1	– 1.775	1
DE	– 6.653*	1941	6		
FL	– 5.122	1967	6	– 1.506	8
GA	– 3.116	1989	1	– 0.887	2
IA	– 7.780*	1948	0		
ID	– 6.078*	1940	8		
IL	– 6.014*	1970	2		
IN	– 5.319	1943	0	– 3.444*	5
KS	– 8.504*	1941	2		
KY	– 4.674	1959	0	– 0.636	7
LA	– 5.182	1985	1	– 2.201	8
MA	– 5.591*	1940	1		
MD	– 5.975*	1942	1		
ME	– 5.529	1943	6	– 1.398	8
MI	– 5.032	1968	3	– 3.865*	5
MN	– 5.685*	1944	3		
MO	– 4.774	1972	2	– 2.058	0
MS	– 5.168	1940	3	– 2.383	8
MT	– 4.850	1951	0	– 3.285	3
NC	– 2.992	1996	8	– 0.995	8
ND	– 6.405*	1950	0		
NE	– 7.263*	1941	6		
NH	– 4.732	1994	8	– 3.939*	8
NJ	– 4.686	1944	1	– 2.416	2
NM	– 4.663	1947	2	– 2.739	2
NV	– 7.135*	1942	4		
NY	– 4.365	1939	1	– 2.818	2
OH	– 5.363	1996	4	– 4.219*	7
OK	– 4.305	1941	8	– 1.526	8
OR	– 4.972	1940	7	– 2.891	0
PA	– 6.951*	1940	0		
RI	– 5.144	1941	8	– 3.458*	8
SC	– 3.479	1991	7	– 2.335	7
SD	– 7.678*	1948	0		
TN	– 4.303	1942	1	– 1.357	2
TX	– 4.275	1941	6	– 2.750	5
UT	– 6.059*	1943	0		
VA	– 5.400	1982	2	– 1.308	5
VT	– 4.830	1946	6	– 3.168	6
WA	– 5.520	1944	1	– 3.728*	1
WI	– 5.333	1963	8	– 2.016	8
WV	– 5.248	1954	3	– 3.921*	1
WY	– 3.791	1991	8	– 2.613	8

* = significant at 5% level (based on critical values from Perron (1997) for endogenous break and MacKinnon (1996) for no break)

he assumed the same fixed break year of 1947 for every state, following from the tests for regional convergence conducted by Carlino and Mills (1993).

A further point should perhaps be made here. Because Perron's test allows for a break in the trend function, additional parameters need to be estimated, and while the test is still valid, its power may be low if in fact no break is called for. On the assumption of no break in the trend function, standard Augmented Dickey-Fuller (ADF) tests can be implemented. The ADF test with constant and trend can be viewed as a restricted form of (6), where $\theta = \gamma = \delta = 0$. Among the states for which the Perron test, with a data-determined break, did not reject the unit-root hypothesis at the 5% level, there were eight for which the unit-root null *can* be rejected at the 5% level on a the basis of the ADF test.[14] We therefore consider these eight additional states to also exhibit stochastic convergence. ADF test statistics for all the states that did not reject the null of non-stationarity under the Perron test are shown in the final two columns of table 4.

3.2. β-convergence

Given these findings, tests for β-convergence are performed on the states that show evidence of stochastic convergence. Any of these states that also support β-convergence would then be classified as converging in income to the national average. Time series tests for β-convergence with an unknown break-point are developed in Vogelsang (2001) and Tomljanovich and Vogelsang (2002). Consider the regression,

$$\dot{y}_t = \mu_1 DU_{1t} + \delta_1 DT_{1t} + \mu_2 DU_{2t} + \delta_2 DT_{2t} + u_t, \tag{7}$$

where $DU_{1t} = 1$ if $t \leq T_B$ and 0 otherwise; $DT_{1t} = t$ if $t \leq T_B$ and 0 otherwise; $DU_{2t} = 1$ if $t > T_B$ and 0 otherwise; $DT_{2t} = t - T_B$ if $t > T_B$ and 0 otherwise; and u_t is the zero mean error term that may be serially correlated. The parameter μ_1 would determine if relative income is above ($\mu_1 > 0$) or below ($\mu_1 < 0$), the average at the start of the series, while δ_1 represents the growth rate until the break year. Likewise, μ_2 would determine if relative income is above or below the average in the estimated break year, and δ_2 represents the growth rate after the break year. The criterion for convergence is based on the coefficients for μ_j and δ_j being of opposite sign, for $j = 1$ before the break year, or $j = 2$ after the break year. This would imply that states with per capita incomes initially above (below) the average grew at a slower (faster) rate than states with per capita incomes initially below (above) the average.

The estimation process is as follows, based on Vogelsang (2001). First, the break year is estimated by choosing the year that maximizes the normalized Wald statistic for no break (i.e., $\mu_1 = \mu_2, \delta_1 = \delta_2$). Trimming is performed to limit the potential break years to lie within the central 80% of the sample period.

14 For this supplementary test we used the same procedure for fixing the lag order, k, as in the Perron test.

Because the break year is unknown, the models are estimated using break years that may differ from the true break years. This results in a misspecification, and thus the standard normal distribution is not applicable. Instead, the critical values presented by Vogelsang (2001) are utilized for standard t-statistics divided by the square root of the sample size. The normalized t-statistics for the nulls of $\mu_j = 0$ and $\delta_j = 0$ are presented in the top panel of table 5 for each state previously found to be stochastically converging. The critical values appear at the bottom of the panel.

Tomljanovich and Vogelsang (2002) perform similar tests for β-convergence on U.S. regional income. Their criterion for evidence of β-convergence in progress is that the estimates for μ_j and δ_j are of opposite sign, and at least one of them is statistically significant. If neither is statistically significant, then β-convergence has already occurred. Divergence would be indicated by t-statistics that are of the same sign, where at least one of them exceeds the critical value (in absolute terms). In addition, those states previously found to be non-stationary (and therefore not included in the β-convergence tests) are also classified as diverging in income.

Based on these criteria, none of the 15 states that we found satisfied stochastic convergence is β-diverging after its break year. In fact, the estimates suggest only DE, IL, and NV are β-converging rather than having already β-converged. Prior to its break year, no state had a significant trend and only CO, IA, ID, MN, and NE were not significantly different from the initial state average, suggesting they had already β-converged. In contrast, DE and NV were initially significantly above the average and remained β-divergent until their break, when they began β-converging. The rest (KS, IL, MA, MD, ND, PA, SD, and UT) were in the process of β-converging and (except for IL) fully β-converged at their break.

In addition, the bottom panel of table 5 shows comparable information for the eight states mentioned earlier, where the Perron test with breaking trend did not reject the unit-root null but a standard ADF test did. For these states we ran β-convergence regressions without a break. The t-statistics shown in the table are of the HAC variety, robust with respect to autocorrelation and heteroscedasticity. The results suggest that each of these states is also converging in income and IN and NH have fully achieved β-convergence.

Thus, we conclude that 23 of the 49 continental states (including DC) are converging or have already converged in income. Therefore, a little less than half of the states exhibit individual evidence of convergence despite the cross-sectional tests failing to reject convergence as a general phenomenon among the states. Note that each of the 26 states that are interpreted as not converging failed to pass the tests for stochastic convergence. Of these, seven are estimated to have break years in the post-1980 period when sigma convergence (and cross-sectional β-convergence) appears to have ended. Of the 40 states where stochastic convergence breaks occurred prior to 1980, a slight majority (21) are found to be converging or fully converged.

TABLE 5
Time series β-convergence

A. Endogenous break

State	μ_1	δ_1	μ_2	δ_2	Break
CO	0.021*	– 0.002	0.049	0.001	1944
	(0.150)	(– 0.173)	(0.745)	(0.503)	
DE	0.456*	0.009	0.339*	– 0.005*	1941
	(2.270)	(0.354)	(4.047)	(– 2.193)	
IA	– 0.205	0.008	0.042	– 0.002	1945
	(– 0.868)	(0.338)	(0.357)	(– 0.539)	
ID	– 0.105	– 0.054	– 0.057	– 0.002	1933
	(– 0.227)	(– 0.384)	(– 0.561)	(– 0.650)	
IL	0.460*	– 0.040	0.290*	– 0.003*	1933
	(2.040)	(– 0.593)	(5.823)	(– 2.551)	
KS	– 0.214*	– 0.004	0.014	– 0.000	1941
	(– 1.226)	(– 0.198)	(0.191)	(– 0.157)	
MA	0.526*	– 0.020	0.102	0.002	1947
	(2.726)	(– 1.209)	(0.981)	(0.778)	
MD	0.296*	– 0.005	0.107	0.002	1944
	(2.284)	(– 0.377)	(1.722)	(0.934)	
MN	– 0.017	– 0.004	– 0.008	0.002	1945
	(– 0.157)	(– 0.352)	(– 0.147)	(1.247)	
ND	– 0.805*	0.030	– 0.096	0.000	1942
	(– 1.571)	(0.498)	(– 0.430)	(0.019)	
NE	– 0.120	– 0.011	0.004	– 0.000	1941
	(– 0.529)	(– 0.397)	(0.045)	(– 0.111)	
NV	0.336*	0.012	0.371*	– 0.005*	1942
	(1.740)	(0.521)	(4.394)	(– 2.473)	
PA	0.252*	– 0.009	0.087	– 0.001	1941
	(3.405)	(– 0.914)	(2.834)	(– 1.299)	
SD	– 0.655*	0.015	– 0.157	0.001	1941
	(– 1.429)	(0.265)	(– 0.824)	(0.226)	
UT	– 0.173*	0.004	– 0.022	– 0.003	1941
	(– 1.041)	(0.167)	(– 0.311)	(– 1.658)	
[cv 95%]	[0.88]	[2.00]	[3.00]	[2.01]	

B. No break

State	μ	δ
AL	– 0.638*	0.007*
	(– 14.591)	(8.632)
IN	0.032	– 0.001
	(0.958)	(– 1.846)
MI	0.232*	– 0.003*
	(9.696)	(– 6.666)
NH	0.015	0.001
	(0.379)	(1.318)
OH	0.196*	– 0.003*
	(19.344)	(– 14.002)
RI	0.268*	– 0.004*
	(5.234)	(– 4.117)
WA	0.189*	– 0.002*
	(8.018)	(– 3.903)
WV	– 0.281*	0.001*
	(– 20.781)	(3.143)

NOTES: Normalized t-statistics in parentheses for Panel A, with presented 95% critical values taken from Vogelsang (2001). Serial correlation-robust t-statistics in parentheses for Panel B. * = significant at 5% level.

3.3. Club convergence

Another possibility is that the states not converging to the *national* average are engaging in club convergence. Clubs can form exogenously at the regional level or endogenously based on initial conditions. Indeed, only two of the 16 southern states were found to be converging, and thus an absolute majority of all the non-converging states (14 of 26) are from this region. Instead, southern states may be forming their own club. As it turns out, all the southern non-converging states were at or below the median income in 1930, and all the non-southern non-converging states were above the median income in 1930.[15] Thus, there is no distinction among the non-converger states between a southern club versus a 'poor' club or a non-southern club versus a 'rich' club.

We investigate club convergence by conducting stochastic and β-convergence tests for each of the non-convergening states, redefining $\dot{y}_t$ in (6) and (7) as relative to the respective 'club' average for the southern/poor or non-southern/rich non-converging states. Results are presented in table 6.[16] Only five southern/poor and three non-southern/rich states are stochastically converging. Among these, Arizona, North Carolina, and South Carolina are β-converging to the southern/poor club, and Florida and Virginia are fully β-converged. Maine and Oregon are fully β-converged to the non-southern/rich club. In constrast, New Jersey is actively β-diverging away from the non-southern/rich club. The results are consistent with one southern/poor club of just five members and one non-southern/rich club of only two members. Thus, there is little evidence of actual club convergence.[17]

4. The determinants of convergence

In this section we attempt to identify characteristics of the individual states that may account for the observation that some states exhibit convergence toward the national average per capita income while others do not. To do so, we construct a binary dependent variable coded as 1 if the state is found to be converging or fully converged in income – which requires that the state pass the tests for both stochastic and β-convergence – and 0 otherwise. We first check to see if certain regions of the country are more prone to convergence than others by regressing the convergence dummy against a set of regional dummies. We construct dummies for Northeast, Midwest, South, and West states based on BEA regions, where

15 Southern states are defined as belonging to the BEA regions of Southeast or Southwest. Arizona was the median income state in 1930, but was closer in income to the next poorer state (Indiana) than to the next wealthier state (Minnesota).

16 To save space, all results are presented in a single table. For simplicity, the number of lags and estimated break years are not reported. For the endogenous break β-convergence tests, only the test statistics for the post-break period are reported.

17 It is possible the 12 non-southern states could be forming separate clubs among their distinctive BEA regions, but with so few potential members in any one club, it further stretches the definition of club convergence.

TABLE 6
Club convergence t-statistics

	Sochastic convergence		β-convergence	
	Perron	ADF	μ	δ
South/poor				
AR	– 5.451	– 2.202		
AZ	– 6.427*		4.210*	– 2.906*
FL	– 6.482*		– 2.605	– 0.305
GA	– 3.514	– 1.851		
KY	– 3.997	– 2.323		
LA	– 4.856	– 3.249		
MS	– 4.763	– 2.902		
NM	– 3.415	– 1.975		
NC	– 5.327	– 3.786*	– 11.098*	7.782*
OK	– 3.641	– 2.214		
SC	– 4.628	– 3.809*	– 11.759*	6.988*
TN	– 4.458	– 3.096		
TX	– 3.967	– 3.144		
VA	– 8.823*		1.488	1.621
Non-South/rich				
CA	– 4.740	– 2.045		
CT	– 4.551	– 2.620		
DC	– 5.308	– 2.178		
ME	– 5.568*		1.385	– 0.180
MO	– 4.510	– 0.741		
MT	– 4.319	– 2.223		
NJ	– 4.984	– 3.579*	8.361*	3.408*
NY	– 4.214	– 2.857		
OR	– 6.833*		0.289	– 1.493
VT	– 4.751	– 2.873		
WI	– 3.932	– 1.373		
WY	– 3.708	– 3.070		

NOTES: β-convergence t-statistics represent post-break for AL, FL, VA, OR, WV. β-convergence tests do not include a break for NC, SC, NJ. * = significant at 5% level.

Northeast includes New England and Mideast regions, Midwest includes Great Lakes and Plains regions, South includes Southeast and Southwest regions, and West includes Rocky Mountain and Far West regions. The Midwest region has the most states that are converging (10/12), followed by the Northeast (6/12), the West (5/9), and finally the South (2/16). The first column of table 7 presents results from a probit regression using the South as the default region.

The regional differences in convergence are statistically significant. A random state from any other region is significantly more likely to be converging compared with a southern state. Because southern states typically represent the poorest states in the nation, our results suggest they are not catching up to the other states. Distinctions among the other regions are less important. The difference between the estimated coefficients for midwestern states and northeast states is of only marginal significance (p-value = .08). Western states are no more or less

TABLE 7
Probit estimates for the probability of state converging

	(I)	(II)	(III)	(IV)	(V)	(VI)	(VII)
Constant	– 1.15*	– 0.44	– 6.61	– 17.24	– 3.79	– 13.61	– 3.48
	(– 2.86)	(– 0.03)	(– 0.43)	(– 1.11)	(– 0.25)	(– 0.88)	(– 0.23)
Midwest	2.12*	2.06*	2.49*	2.44*	2.33*	2.44*	2.54*
	(3.60)	(2.61)	(2.85)	(2.23)	(2.97)	(2.63)	(2.74)
Northeast	1.15*	0.87	1.06	0.88	0.94	1.02	0.97
	(2.13)	(1.20)	(1.44)	(1.24)	(1.30)	(1.42)	(1.32)
West	1.29*	1.09	1.39	1.03	1.21	1.17	1.47
	(2.22)	(1.26)	(1.52)	(0.94)	(1.45)	(1.26)	(1.48)
K/L Ratio Change		– 0.36*	– 0.48*	– 0.37*	– 0.39*	– 0.47*	– 0.40*
		(– 2.34)	(– 2.67)	(– 2.07)	(– 2.07)	(– 2.72)	(– 2.25)
Patents Stock		0.04	– 0.26	0.09	– 0.17	– 0.12	– 0.15
		(0.09)	(– 0.49)	(0.17)	(– 0.32)	(– 0.24)	(– 0.30)
Education		0.64	0.46	3.43	0.55	2.04	0.21
		(0.17)	(0.13)	(0.89)	(0.15)	(0.54)	(0.06)
Agricultural Sector Share		– 0.40	– 0.75*	– 0.85*	– 0.63*	– 0.81*	– 0.72*
		(– 1.60)	(– 2.14)	(– 2.19)	(– 1.92)	(– 1.94)	(– 2.03)
Government Share		– 0.66					
		(– 0.89)					
Economic Freedom Index			0.78				
			(1.43)				
EFI Category 1				– 0.48	0.26		
(government spending)				(– 0.73)	(0.68)		
EFI Category 2				1.38*		1.00*	
(tax rates and revenue)				(1.96)		(1.94)	
EFI Category 3				– 0.10			0.47
(labour market freedom)				(– 0.12)			(0.83)
Regional Effect	4.44	2.36	2.86	2.40	3.08	2.57	2.74
(*p*–value)	(0.01)	(0.09)	(0.05)	(0.08)	(0.04)	(0.06)	(0.05)
n	49	49	48	48	48	48	48
Pseudo R^2	0.23	0.32	0.34	0.38	0.32	0.37	0.33

NOTES: *z*-statistics derived from QML standard errors in parentheses. Economic Freedom Index is the average of the three EFI categories, where higher values for each category represents more economic freedom (less government spending, taxes and/or labour regulations). * = significant at 5% level.

likely to converge than the northeastern states (equality of coefficients can be rejected only with *p*-value = .80) or midwestern states (*p*-value = .17).

We next seek to determine if we can exlain the regional differences that do exist. According to neoclassical growth theory, economic growth is determined by capital, labour, and new technology, and possibly is augmented by human capital and other factors (Barro and Sala-i-Martin 1991). In the absence of mobility barriers, resources are expected to flow to where they are most productive. As workers move to higher income areas, the capital to labour (K/L) ratio will fall in wealthier areas, depressing wages and slowing growth. This process allows for catch-up. Mobility that affects the capital to labour ratio is therefore expected to be a primary determinant of convergence. If workers are attracted to high-wage areas, then inter-state labour force migration should reduce the K/L ratio

in high-income areas and increase it in low-income areas. International labour force migration should flow more heavily into high-income areas, reducing the K/L ratio to a larger extent in high-income areas. Combining these two phenomena, high-income areas should have their ratios affected to a larger extent than low-income areas. We therefore expect convergence to be *inversely* related to the change in the K/L ratio. Our measure of capital is based on the state-level net private capital stock series developed by Garofalo and Yamarik (2002), which begins in 1947. We therefore consider the change in the capital to labour ratio from 1950 to 2000, where the denominator is measured by the total civilian labour force.[18]

Technology is proxied by the number of patents per capita compiled by Hunt and Gauthier-Loiselle (2010). They collected data in 10-year intervals beginning from 1930 to 2000 (and also 2005). To match the K/L ratio change data range, we use the log of the average value of patents per capita from 1950 to 2000.[19] The effect of patents, though, on convergence is ambiguous. Technology can be considered a positive shock to growth. As a proxy for technology, patents should help poor states to catch up, thereby assisting in the β-convergence process. Yet patents prevent the diffusion of new technology, [20] enabling the possessor to reap the rewards for a longer period. Although this will help with 'catch-up', it also suggests the impact from a temporary technology shock will last much longer than if the technology can be quickly copied by others. Thus, a state experiencing a one-time positive shock to its relative income will not have a tendency to quickly return to its previous trend, implying a unit root (or near unit root). Patents, therefore, limit the potential for a state to experience stochastic convergence. Because patents are expected to be positively correlated with the likelihood of β-convergence but inversely correlated with the likelihood of stochastic convergence, the net effect on overall convergence is ambiguous.

If technology is not patented (or once the patent expires), the ability to adopt new technologies depends on the skills of the labour force. Greater human capital should therefore be positively correlated with convergence. We proxy the level of human capital during this period by the log of the average value from 1950 to 2000 for the percentage of the population having completed high school. Data for each decadal value are taken from the Census website.

We supplement these determinants with two measures for the sectoral composition of output taken from BEA. The extent of economic development is proxied by the percentage of state GSP from agriculture. States that are more heavily dependent on farm income are typically less developed than those engaged in manufacturing and services. Agricultural sectors are also more prone to

18 Reported results are very similar if we use the employed labour force instead of total labour force.

19 Note also, that according to table 5, several states appear to have had their β-convergence break in the 1940s. β-convergence for these states is strictly identified in the post-break period.

20 We note that patented technology is not limited to the state where the patent is held. For example, a company headquartered in one state can freely adopt its patented technology for any of its plants, even if they are located in different states. Still, the number of other states affected by a patent should be small relative to the number of states in the country.

experiencing, and less able to adjust to, negative shocks. Thus, states more heavily dependent on agricultural output should be less likely to be converging. We measure agricultural output by the the percentage of total income represented by farm income, averaged from 1950–2000, then logged.

Finally, one purpose behind redistributive government is to help weather shocks. A more activist government sector engaging in counter-cyclical policies should enable a state to smooth its business cycles, helping the adjustment process to various shocks, tempering their impact. Thus, stochastic convergence should be more likely to occur. The size of the government sector is measured similarly to that of the agricultural sector: the log of the average decade values from 1950 to 2000 for the percentage of total state income coming from the government sector.

As shown in column (II), adding the specific controls described above eliminates significance of the individual regional variables for Northeast and West, but not for Midwest. Furthermore, although equality between coefficents for the West and either of the other regional coefficients cannot be rejected at conventional levels, states from the Midwest remain statistically signficantly more likely to converge compared to southern states at better than the 1% level. Thus, some extent of regional deviations in convergence remain unexplained even after including these additional variables. Among the added controls, most coefficients are of the expected sign, but only the K/L ratio change is found to be statistically significant at better than the 5% level. Perhaps surprisingly, the government share of output reduces the likelihood of convergence, albeit not to a statistically significant degree.

Although the government variable directly represents sectoral composition, the size of government has also been sometimes used by growth scholars to proxy for the institutional environment (see the recent surveys by Holcombe 2001 and Mueller 2003). In particular, government spending has been interpreted by many to represent market distortions. Under this interpretation, the negative coefficient for government size makes more sense. To better determine if market distortions are indeed making convergence more difficult, a more direct measure for the institutional environment / market distortions would be useful. Recently, a set of economic freedom measures for the U.S. states has been developed by the Fraser Institute to mimic their more comprehensive national economic freedom indices, which have been widely used in various growth studies.[21] Government consumption is but one of several components to the state economic freedom index. The index rates the states on a scale from 0 to 10 with higher scores representing greater levels of economic freedom (less government intervention and regulation) in the areas of Size of Government, Takings and Discriminatory Taxes, and Labour Market Freedom (Ashby et al. 2010).[22]

21 De Haan, Lundström, and Sturm (2006) represents a recent survey of the vast economic freedom literature.

22 The Size of Government score is based on the average value of component scores reflecting government consumption, transfers and subsidies, and social security payments. The Takings

The state-level economic freedom index (EFI) is not available until 1981, and then only for actual states, not for District of Columbia. We use the average values for 1981–2000 to generate an EFI score for each state and use it to replace the government share variable. As shown in column (III) of table 7, the EFI is not significantly correlated with the likelihood of convergence. The positive coefficient suggests that more freedom from government increases the probability of convergence, although not to a statistically significant degree. This result is consistent with the negative coefficient on the government sector share reported in the previous column.[23] None of the other variable coefficients is greatly affected, except that agriculture share is now signficant at better than 5%, and the regional effect is somewhat strengthened.

Breaking the EFI into its three component ratings in column (IV) reveals an important distinction among the institutional/policy indicators. Components 1 (government spending) and 3 (labour market freedom) generate negative but statistically insignificant coefficients. They are also not jointly significant (p-value = .72). Rather, the extent of tax rates and revenue, as captured in component 2, is statistically significant. Because higher values represent less intervention, the positive coefficient indicates that higher taxes generally hinders the prospect for convergence.

The EFI categories are moderately correlated, ranging from 0.43 (categories 2 and 3) to 0.58 (categories 1 and 3). The last three columns include each of the EFI categories in its own regression. The results are similar. The coefficient for EFI category 2 remains positive and statistically significant. Coefficients for EFI categories 1 and 3 change sign from negative to positive, but remain far from being statistically significant, again suggesting no independent effect. Thus, the impact, or lack of impact, of each EFI category reported in column (IV) was not dependent on multicollinearity among the types of freedom included in the EFI.

In sum, convergence appears to be a function of changing capital to labour ratios, as increases in the labour force relative to capital (presumably representing greater labour migration into high-income areas) benefits convergence. States more heavily dependent on agriculture and taxes are less likely to be converging. There are also regional differences in the likelihood of state convergence, especially among the midwestern (more likely) and southern (less likely) states. Thus, assuming that southern states are unable to physically relocate to the midwest (!), policymakers can assist in the convergence process by committing to low taxes and offering incentives for conversion away from agricultural production, perhaps by reducing agricultural subsidies and other forms of protectionism. Lowering barriers to international immigration (mostly, but not exclusively, a

and Discriminatory Taxes score represents the average score for total tax revenues, top marginal tax rates, indirect tax revenues, and sales tax revenue. Labour Market Freedom is determined by the average score on minimum wage laws, government employment, and union density.

23 Using the log of government share average from 1980 to 2000 to match the EFI years still results in a negative but statistically insignificant coefficient.

federal issue; intranational migration is limited almost exclusively by transaction costs of moving) might also assist in convergence.

5. Conclusions

We perform per capita income convergence tests on the U.S. states (including District of Columbia). Cross-sectional tests support σ-convergence (reduced dispersion) and β-convergence (poorer states grow faster than wealthier states) over the full 80-year period, 1930–2009, but these results may not hold true for the last three decades. This supports and extends the findings of Crown and Wheat (1995) and Barro and Sala-i-Martin (1991).

For individual states to be considered converging, they must be both 'stochastically converging' – implying stability or reversion to trend after shocks – and β-converging. Time series tests covering annual data over this period suggest that only 23 of the 49 continental states exhibit stochastic convergence. Of these, all are also in the process of or have already achieved β-convergence and thus would be considered to be converging in income.

Probit regressions reveal that midwestern states are more likely to exhibit convergence than are southern states. In addition, states with smaller government sectors (especially in regard to taxation), less reliance on agricultural output, and greater changes in the capital to labour ratio through attracting a greater labour force, are significantly more likely at the margin to exhibit convergence. The extent of human capital and patent assignments in the state, however, does not appear to affect the likelihood of convergence at the margin. These findings suggest that poorer states, especially in the south, should focus on economic development through sectoral transition to manufacturing and services, and at the same time reduce taxes. Offering tax concessions on firm relocation and production might be one way to accomplish both ends.

These results on convergence are limited to nominal state incomes. It is possible that additional states could be converging in real income if their price dynamics differ substantially from other states. It might therefore be useful to supplement these results by testing for real income convergence. Unfortunately, state price indexes are not available prior to 1977, and alternative time-series testing techniques might be needed for such a short time span with many fewer observations.

References

Ashby, Nathan J., Amela Karabegović, Fred McMahon, and Avilia Bueno (2010) *Economic Freedom of North America 2010*. Vancouver: Fraser Institute

Barro, Robert J., and Xavier Sala-i-Martin (1991) 'Convergence across states and regions.' *Brookings Papers on Economic Activity* 1, 107–58

— (1992) 'Convergence.' *Journal of Political Economy* 100, 223–51

Carlino, Gerald A., and Leonard O. Mills (1993) 'Are U.S. regional incomes converging?' *Journal of Monetary Economics* 32, 335–46
— (1996) 'Testing neoclassical convergence in regional incomes and earnings.' *Regional Science and Urban Economics* 26, 565–90
Choi, Chi-Young (2004) 'A reexamination of output convergence in the U.S. states: toward which level(s) are they converging?' *Journal of Regional Science* 44, 713–41
Conlan, Timothy J., and Paul L. Posner (2012) 'Inflection point? Federalism and the Obama administration.' *Publius* 41, 421–46
Coughlin, Cletus C., Thomas A. Garrett, and Rubén Hernández-Murillo (2004) 'Spatial probit and the geographic patterns of state lotteries.' Working Paper 2003-042B, Federal Reserve Bank of St. Louis
Crown, William H., and Leonard F. Wheat (1995). 'State per capita income convergence since 1950: Sharecropping's demise and other influences.' *Journal of Regional Science* 35, 527–52
De Haan, Jakob, Susanna Lundström, and Jan-Egbert Sturm (2006). 'Market oriented institutions and policies and economic growth: A critical survey.' *Journal of Economic Surveys* 20, 157–91
Evans, Paul (1997) 'How fast do economies converge?' *Review of Economics and Statistics* 79, 219–25
Fingleton, Bernarnd, and Enrique López-Bazo, (2006) 'Empirical growth models with spatial effects.' *Papers in Regional Science* 85, 177–98
Gaffikin, Frank, and Barney Warf (1993) 'Urban policy and the post-Keynesian state in the United Kingdom and the United States.' *International Journal of Urban and Regional Research* 17, 67–84
Garofalo, Gasper A., and Steven Yamarik (2002) 'Regional growth: evidence from a new state-by-state capital stock series.' *Review of Economics and Statistics* 84, 316–23
Goetz, Stephan J., and Dayuan Hu (1996) 'Economic growth and human capital accumulation: Simultaneity and expanded convergence tests.' *Economics Letters* 51, 355–62
Greene, William H. (2008) *Econometric Analysis*. 6th ed. Upper Saddle River, NJ: Pearson Prentice Hall
Holcombe, Randall (2001) 'Public choice and economic growth.' In *The Elgar Companion to Public Choice*, ed. W.F. Shughart II and L. Razzolini. Cheltenham, UK: Edward Elgar Press
Hunt, Jennifer, and Marjolaine Gauthier-Loiselle (2010) 'How much does immigration boost innovation?' *American Economic Journal: Macroeconomics* 2, 31–56
Izraeli, Oded, and Keven J. Murphy (1997) 'Convergence in state nominal and real per capita income.' *Public Finance Review* 25, 555–76
Johnson, Paul A. (2000) 'A nonparametric analysis of income convergence across the U.S. states.' *Economics Letters* 69, 219–23
Kenworthy, Lane (1999) 'Economic integration and convergence: a look at the U.S. states.' *Social Science Quarterly* 80, 858–69
MacKinnon, James G. (1996) 'Numerical distribution functions for unit root and cointegration tests.' *Journal of Applied Econometrics* 11, 601–18
Maddala, G.S., and Shaowen Wu (2000) 'Cross-country growth regressions: problems of heterogeneity, stability and interpretation.' *Applied Economics* 32, 635–42
Miller, Stephen M., and Frank S. Russek (1997) 'Fiscal structures and economic growth at the state and local level.' *Public Finance Review* 25, 213–37
Mueller, Dennis C. (2003) *Public Choice III*. Cambridge: Cambridge University Press
Olson, Mancur (1982) *The Rise and Decline of Nations*. New Haven, CT: Yale University Press
— (1983) 'The south will fall again: the south as leader and laggard in economic growth.' *Southern Economic Journal* 49, 917–32

Owyang, Michael T., Jeremy Piger, and Howard J. Wall (2008) 'A state-level analysis of the great moderation.' *Regional Science and Urban Economics* 38, 578–89

Perron, Pierre (1997) 'Further evidence on breaking trend functions in macroeconomic variables.' *Journal of Econometrics* 80, 355–85

Quah, Danny T. (1993a) 'Empirical cross-section dynamics in economic growth.' *European Economic Review* 37, 426–34

— (1993b) 'Galton's fallacy and tests of the convergence hypothesis.' *Scandinavian Journal of Economics* 95, 427–43

— (1996) 'Empirics for economic growth and convergence.' *European Economic Review* 40, 1353–75

Rey, Sergio J., and Brett D. Montouri (1999) 'U.S. regional income convergence: a spatial econometric perspective.' *Regional Studies* 33, 143–56

Tomljanovich, Marc, and Timothy J. Vogelsang (2002) 'Are U.S. regions converging? Using new econometric methods to examine old issues.' *Empirical Economics* 27, 49–62

Vogelsang, Timothy J. (2001) 'Testing for a shift in trend when serial correlation is of unknown form.' Working Paper No. 97–11, Center for Analytic Economics, revised October