ELSEVIER Journal of Monetary Economics 38 (1996) 587–598

JOURNAL OF Monetary ECONOMICS

Are U.S. regional incomes converging? Some further evidence

Michael B. Loewy*, David H. Papell

Department of Economics, University of Houston, Houston, TX 77204-5882, USA

(Received August 1994; final version received February 1996)

Abstract

Regions of the United States represent a set of economies where the conditions often argued as underlying convergence in per-capita income are satisfied. Stochastic convergence among regions implies the rejections of a unit root in relative per-capita income, a hypothesis which is not supported by conventional tests. Carlino and Mills (1993) allow for an exogenous trend break, but can only reject the unit root hypothesis in three of the eight U.S. regions. We incorporate endogenously determined break points and significantly strengthen their results, being able to reject the unit root hypothesis in seven regions.

Key words: Regional per-capita income; Stochastic convergence; Endogenous trend break models

JEL classification: C32; O40

1. Introduction

One of the most noteworthy implications of Solow's (1956) neoclassical growth model is that once the determinants of steady-state per-capita income have been controlled for, economies exhibit convergence. In particular, at a point in time there exists a negative relationship between initial log per-capita income and rates of growth. Given its dependence upon the factors determining

*Corresponding author.

We wish to thank Dan Ben-David, Gerald Carlino, Steven Craig, Janet Kohlhase, the editor, and the referee for supplying us with useful comments. We also wish to thank Gerald Carlino for supplying us with the data. Any remaining errors are the sole responsibility of the authors.

PII S0304-3932(96)01292-5

the steady state, Mankiw, Romer, and Weil (1992) and Barro and Sala-i-Martin (1992) refer to this form of convergence as *conditional convergence*.

In contrast to this cross-section notion of convergence, the existence of random, but potentially permanent, shocks to per-capita income have lead researchers to formulate a time-series notion of convergence. Assuming that an economy's log per capita income possesses a unit root, Campbell and Mankiw (1989) and Bernard and Durlauf (1995) define *stochastic convergence* as cointegration between two (or more) such series. An alternative definition of stochastic convergence, due to Carlino and Mills (1993), is that the log of the percapita income of one region relative to that of the economy as a whole is stationary.[1]

Recently, interest in economic growth and the question of convergence has been renewed in part by the critiques of the neoclassical growth model offered by Romer (1986) and Lucas (1988), among others. These criticisms have lead to a large literature dedicated to applying cross-section and time-series techniques to both cross-country and cross-regional data in order to determine whether such economies exhibit conditional and/or stochastic convergence.

In the cross-section setting, Barro (1991) and Mankiw, Romer, and Weil (1992) provide evidence showing that while conditional convergence does not hold across most countries as a whole, it does arise among groups of countries with certain characteristics in common. In contrast, the time-series evidence suggests that convergence fails to hold. Quah (1990) finds little evidence of cross-country stochastic convergence among a large set of capitalist economies while Campbell and Mankiw (1989) and Bernard and Durlauf (1995) find likewise for sets of OECD economies.

Turning to cross-regional data, Barro and Sala-i-Martin (1991, 1992) find much cross-section support for conditional convergence across both regions of Western Europe and U.S. states. On the other hand, Brown, Coulson, and Engle (1990) find little time-series evidence supporting cointegration, and hence of stochastic convergence, across a number of U.S. states.

Though the inconsistency between the cross-section and the time-series literatures on cross-country convergence among similar economies is surprising, that it should also be present in the cross-regional data for the U.S. is even more unexpected. The U.S. states represent a set of economies in which there exists nearly complete free trade and mobility of factors, and nearly identical forms of government so that many of the conditions often argued as underlying convergence obtain most strongly. In contrast, OECD countries, though similar in many ways, clearly satisfy these conditions to a lesser degree.[2]

---

[1] For a useful discussion of the relationships between the cross-section and time-series notions of convergence, see Bernard and Durlauf (1996).

[2] For further discussion, see Barro (1991) and Ben-David (1994).

Carlino and Mills (1993), henceforth CM, have recently addressed this inconsistency by investigating both stochastic and conditional convergence (what they refer to as β-convergence) for the eight U.S. regions defined by the Bureau of Economic Analysis. Their initial results regarding stochastic convergence are not encouraging. Using conventional tests, they cannot reject the unit root hypothesis in the log of relative (to the U.S. as a whole) regional per-capita income for any region.[3] Utilizing the techniques of Perron (1989), CM proceed to allow for the existence of an exogenously imposed trend break. After doing so, they report evidence rejecting the unit root (and hence supporting stochastic convergence) in three out of eight regions and conditional convergence in all regions. Thus, by allowing for a trend break, CM are able to eliminate some of the inconsistency between the cross-section and time-series evidence on U.S. regional convergence.[4]

In this paper, we investigate further the purported inconsistency between the cross-section and time-series evidence on convergence among U.S. regions. Rather than assuming an exogenously imposed trend break as CM do, we let 'the data speak for itself' by allowing for endogenously determined break points. We estimate both additive outlier models in which the break occurs instantaneously and innovative outlier models in which the break occurs slowly over time. By so doing, we significantly strengthen CM's results, being able to reject the unit root hypothesis in seven out of eight regions for at least one of the two models. This in turn implies that regional incomes in the U.S. are indeed stochastically converging to that of the nation as a whole as predicted by the neoclassical growth model.

The rest of this paper is organized as follows. Section 2 discusses endogenous versus exogenous trend breaks and provides our empirical results. Section 3 concludes.

2. Endogenous vs. exogenous trend breaks

To provide a frame of reference for what follows, Fig. 1 contains plots of the log relative per-capita income for each of the eight BEA regions. Not surprisingly, the Northeast, Mideast, Great Lakes, and Far West regions typically have per-capita incomes above the national average, while the Plains, Southeast, Southwest, and Rocky Mountains regions typically lie below the national average. Of interest here is whether any of these regions are stochastically converging, that is, have a log relative per-capita income that is stationary about zero.

[3]This conclusion is further supported by various test for the presence of persistence in the data.

[4]CM also report the presence of little or no persistence once the trend break is imposed.

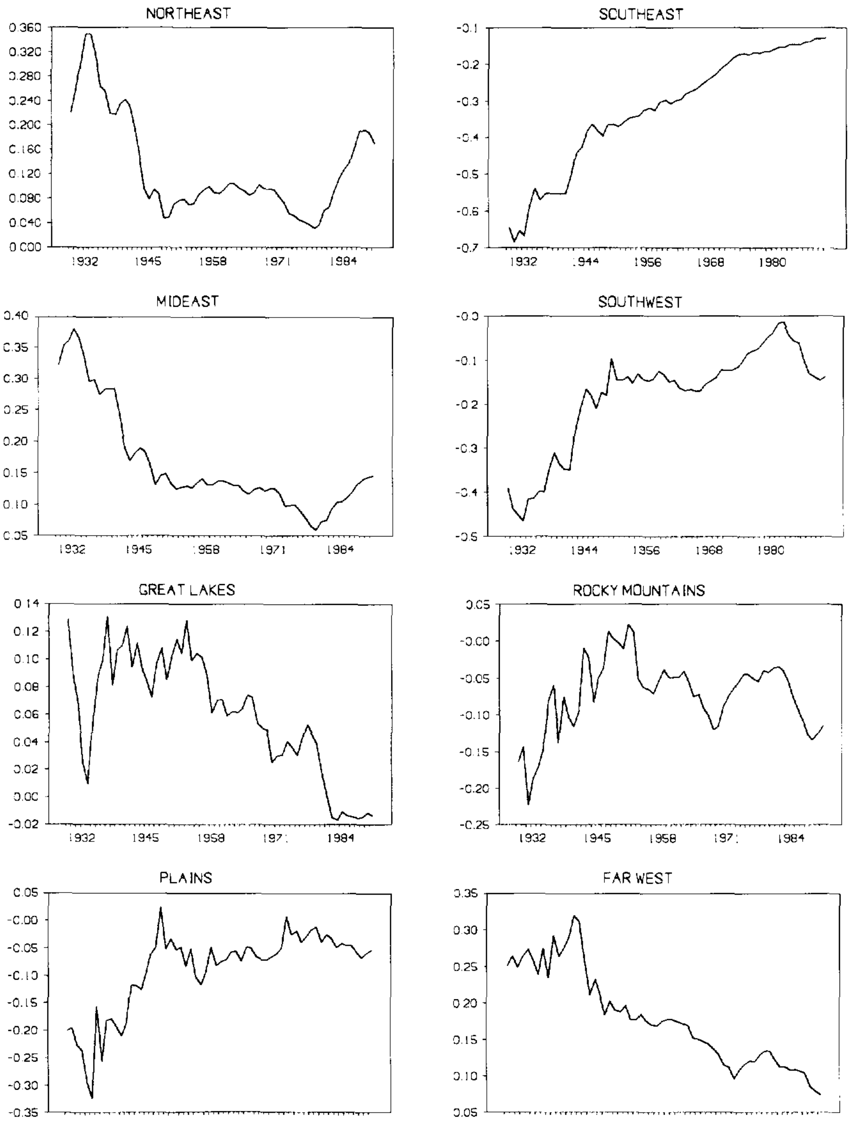

Fig. 1

In the relative per-capita income formulation used both by CM and by us, the existence of stochastic convergence is related to the unit root hypothesis. If shocks to relative regional per-capita incomes are temporary, then relative income will not have a unit root. Using Augmented Dickey–Fuller (ADF) tests without trend breaks, CM are unable to reject the unit root hypothesis for eight U.S. regions during the 1929–1990 period. In this context, failure to reject the unit root null corresponds to failure to provide evidence of stochastic convergence.

In order to provide a benchmark for our results, we first investigate unit roots without trend breaks. Using the same data as CM, regional per-capita income relative to national per-capita income for the U.S. during 1929–1990, we also estimate ADF tests.[5] Since the data are clearly trending, we use the form of the ADF test with both drift and trend,

$$\Delta RI_t = \mu + \beta t + \alpha RI_{t-1} + \sum_{j=1}^{k} c_j \Delta RI_{t-j} + \varepsilon_t, \tag{1}$$

where RI_t is the log of relative regional per-capita income at time t.[6]

Although CM set the lag length, k, equal to one, considerable evidence exists that data-dependent methods to select the value of k are superior to choosing a fixed k a priori. We follow the procedure, denoted $k = \kappa(t\text{-stat})$, suggested by both Campbell and Perron (1991) and Ng and Perron (1995). Start with an upper bound, $k_{\max}$, on k. If the last included lag is significant, choose $k = k_{\max}$. If not, reduce k by one until the last lag becomes significant. If no lags are significant, set $k = 0$. Following Perron (1989), we set $k_{\max} = 8$ and use the (approximate) 10% value of the asymptotic normal distribution, 1.60, to assess the significance of the last lag.[7]

The results of the ADF tests are reported in the top panel of Table 1. In order to provide comparability with the results of CM, we report results both when $k = 1$ and when $k = \kappa(t\text{-stat})$. The null hypothesis of a unit root is rejected if the t-statistic for α is greater (in absolute value) than the appropriate critical value. While the critical values (which appear in the bottom panel of Table 1) are nonstandard, they are widely available (a recent source is MacKinnon, 1991). When $k = 1$, the unit root hypothesis cannot be rejected at the 10% level for any region. This is the same result found by CM. When $k = \kappa(t\text{-stat})$, the unit root null cannot be rejected at the 5% level for any series, but can be rejected at the 10% level for the Far West region.

---

[5]As CM (1993, Fn. 4) point out, though it is preferable to deflate per-capita incomes, the absence of such data for regions makes doing so impossible. Furthermore, doing so using the national CPI is unwarranted since we are considering relative per-capita incomes.

[6]In order to discuss conditional convergence, CM assume that the underlying error process for RI is ARMA(2,0). This restriction does not affect the unit root results.

[7]Ng and Perron (1995) use simulations to show that these sequential tests have an advantage over information-based methods since the former produces tests with more robust size properties without much loss of power.

Table 1
Augmented Dickey–Fuller tests

	$k = 1$		$k = \kappa$ (t-stat)		
Region	α	t_α	α	t_α	k
New England	− 0.07	− 2.30	− 0.07	− 2.30	1
Mideast	− 0.06	− 1.56	− 0.07	− 1.49	6
Great Lakes	− 0.22	− 2.60	− 0.27	− 3.12	2
Plains	− 0.17	− 1.80	− 0.17	− 1.80	1
Southeast	− 0.13	− 2.44	− 0.17	− 2.98	8
Southwest	− 0.08	− 1.42	− 0.13	− 2.22	5
Rocky Mountains	− 0.18	− 2.24	− 0.19	− 2.52	0
Far West	− 0.26	− 2.57	− 0.30	− 3.29*	0
Tests with an exogenous trend break in 1946					
New England	− 0.23	− 4.39**	− 0.23	− 4.39**	2
Mideast	− 0.29	− 3.42	− 0.22	− 1.52	6
Great Lakes	− 0.50	− 4.13*	− 0.80	− 3.93*	2
Plains	− 0.63	− 4.17**	− 1.08	− 4.15*	8
Southeast	− 0.36	− 3.19	− 0.39	− 2.51	8
Southwest	− 0.28	− 2.89	− 0.34	− 2.38	5
Rocky Mountains	− 0.46	− 3.78	− 0.41	− 3.85	0
Far West	− 0.50	− 3.10	− 0.48	− 3.94*	0

Critical values

	1 percent	5 percent	10 percent
ADF test	− 3.96	− 3.43	− 3.13
Exogenous break test	− 4.78	− 4.17	− 3.87

***, **, and * denote statistical significance at the 1%, 5%, and 10% levels, respectively.

It is well-known, however, that the failure of ADF tests to reject the unit root hypothesis may be caused by misspecification of the deterministic trend (Campbell and Perron, 1991). Following Perron (1989), CM allow for a one-time break in the trend function which occurs in 1946. Once this break is imposed, they can reject the unit root null at the 5% level for two regions, Great Lakes and Plains, and at the 10% level for Rocky Mountains. They interpret these results as evidence of stochastic convergence.

There are two types of trend break models: the additive outlier (AO) model where the break occurs instantaneously and the innovational outlier (IO) model where the break is assumed to follow the same path as the innovations and occurs slowly over time. Although CM only estimate the IO model, we see no compelling reason to specify the form of the break ex ante, and estimate both.

The IO model involves estimating the following regression:

$$\Delta RI_t = \mu + \beta t + \delta D(T_B)_t + \theta DU_t + \gamma DT_t + \alpha RI_{t-1} + \sum_{j=1}^{k} c_j \Delta RI_{t-j} + \varepsilon_t, \tag{2}$$

where, denoting the break date as T_B, $D(T_B)_t = 1$ if $t = T_B + 1$, 0 otherwise; $DU_t = 1$ if $t > T_B$, 0 otherwise; and $DT_t = t - T_B$ if $t > T_B$, 0 otherwise. The unit root null is rejected if the t-statistic for α is sufficiently large (in absolute value).

We first estimate the IO model with the exogenous break date set at 1946. The estimates in CM differ from Eq. (2) in two respects, the first being the lag length criteria described above. The second is that CM do not precisely follow Perron's specification for the dummy variables. Instead, they include four dummies: $D_1 = 1$ if $t < T_B$, 0 otherwise; $D_2 = 0$ if $t < T_B$, 1 otherwise; $DT_1 = t$ if $t < T_B$, 0 otherwise; and $DT_2 = 0$ if $t < T_B$, $t - T_B$ otherwise. As in the ADF test, the unit root null is rejected if the t-statistic for α is greater (in absolute value) than the appropriate critical value. The critical values, which are greater (in absolute value) than those for the ADF test, are reported by Perron (1989).

Estimates of the IO model with the break date set at 1946 are reported in the middle panel of Table 1, while the critical values again appear in the bottom panel of the table. When $k = 1$, the unit root null is rejected at the 5% level for New England and Plains and at the 10% level for Great Lakes. These results are similar to those of CM except for New England, where they could not reject the unit root null at 10%. When $k = \kappa(t\text{-stat})$, the results accord with the model with $k = 1$ except for Far West, for which the null can be rejected at 10%.[8]

The testing procedure followed by Perron and CM, where the break date is chosen exogenously, has been the subject of much criticism. Since the choice of break points are based on prior observation of the data, there are obvious problems of pre-testing associated with the methodology. In response to these problems, Banerjee, Lumsdaine, and Stock (1992), Christiano (1992), and Zivot and Andrews (1992), among others, have developed tests where the break is data-dependent. These tests often find less evidence against the unit root hypothesis than do exogenous break tests. In particular, Zivot and Andrews cannot reject the unit root hypothesis at the 5% level for four of the ten Nelson and Plosser (1982) series for which Perron rejects the hypothesis.

The IO model with an endogenous trend break is implemented by estimating Eq. (2) sequentially for each break year $T_B = k + 2, \ldots, T - 1$, where T is the number of observations. The break year that is chosen is the one which

[8]Estimates using CM's dummy variables (not reported) verify that the differences between our results and theirs is due to the different dummy variables.

Table 2
Sequential IO trend break model

	$k = 1$			$k = \kappa(t\text{-stat})$			
Region	Break year	α	t_α	Break year	α	t_α	k
New England	1940	− 0.20	− 4.53	1940	− 0.20	− 4.53	1
Mideast	1939	− 0.25	− 4.31	1948	− 0.52	− 4.01	8
Great Lakes	1932	− 0.39	− 5.31**	1951	− 1.61	− 5.32**	8
Plains	1950	− 0.71	− 4.92*	1949	− 0.69	− 5.78***	0
Southeast	1939	− 0.38	− 5.07*	1940	− 0.51	− 6.78***	3
Southwest	1940	− 0.32	− 4.12	1940	− 0.36	− 4.08	8
R. Mountains	1951	− 0.56	− 4.45	1951	− 0.49	− 4.44	0
Far West	1942	− 0.55	− 5.75***	1942	− 0.65	− 5.39**	7

Critical values

1 percent	5 percent	10 percent
− 5.57	− 5.08	− 4.82

***, **, and * denote statistical significance at the 1%, 5%, and 10% levels, respectively.

minimizes the t-statistic for α. The lag length is chosen both by setting $k = 1$ and using the $k = \kappa(t\text{-stat})$ data-dependent method described above. In order to provide a comparison between our results and those of CM, we use the asymptotic critical values, which are larger (in absolute value) than those in Perron (1989), provided in Zivot and Andrews (1992) and Perron (1994).[9]

The results for the endogenous trend break IO model are presented in Table 2. With $k = 1$, the unit root null can be rejected at the 1% level for Far West, the 5% level for Great Lakes, and the 10% level for Plains and Southeast. With $k = \kappa(t\text{-stat})$, the null can be rejected at 1% for Plains and Southeast and at 5% for Great Lakes and Far West. Despite the larger critical values, these results more strongly support stochastic convergence than do those of CM who can only reject the unit root null at the 5% level for Great Lakes and Plains. The evidence for stochastic convergence is stronger when the lag length, as well as the break year, is chosen endogenously. The break years cluster into two groups; one at the beginning of World War II and a second in 1948–1951. None of the regions has a break in 1946.

[9] Although Zivot and Andrews (1992) do not incorporate the dummy variable $D(T_B)_t$, this does not affect the asymptotic critical values.

The AO model is estimated by a two-step procedure. The first step involves estimating the following regression:

$$RI_t = \mu + \beta t + \theta DU_t + \gamma DT_t + \rho_t, \tag{3}$$

where ρ_t is the residual; and, as above, $DU_t = 1$ if $t > T_B$, 0 otherwise; and $DT_t = t - T_B$ if $t > T_B$, 0 otherwise. The second step involves testing the null hypothesis that $\alpha = 0$ in the regression:

$$\Delta\rho_t = \sum_{i=0}^{k} \omega_i D(T_B)_{t-i} + \alpha\rho_{t-1} + \sum_{j=1}^{k} c_j \Delta\rho_{t-j} + \varepsilon_t, \tag{4}$$

where $D(T_B)_{t-i} = 1$ if $t = T_B - i + 1$, 0 otherwise. The dummy variables are needed to ensure that the t-statistic on α in Eq. (4) has the same asymptotic distribution as in the IO model and is invariant to the value of k. Since CM do not estimate an AO model, we only consider the endogenous trend break version. Eqs. (3) and (4) are estimated sequentially for each possible break, and the choice of break year and lag length are based on the same criteria as for the IO model described above. The asymptotic critical values are the same as in the IO model. Vogelsang and Perron (1994) provide an extensive discussion of the AO model, including additional methods for choosing the break year.[10]

The results for the endogenous trend break AO model are presented in Table 3. With $k = 1$, the unit root null can be rejected at the 5% level for Southeast and Far West and at the 10% level for New England and Plains. With $k = \kappa(t\text{-stat})$, the null can be rejected at 1% for Plains, at 5% for Great Lakes, Southwest, and Far West, and at 10% for Mideast. There is again a cluster of regions with breaks at the beginning of World War II, and no breaks in 1946. As with the IO model, the endogenous break AO model provides more evidence against the unit root hypothesis than is found by CM, and this evidence is stronger when k is data dependent than when k is fixed.

Based on the Zivot and Andrews (1992) results, one might have expected the endogenous tests to provide less evidence against the unit root null in relative regional per-capita incomes than do the exogenous tests. However, this is not the case. We can reject the unit root hypothesis at the 1% level for three of the series, at the 5% level for two more series, and at the 10% level for two additional series for at least one of the two models. These results provide much more evidence against the unit root hypothesis, and thus support for stochastic convergence,

[10] These methods involve maximizing (or minimizing) the t- or F-statistics for the dummy variables. We tried this for both the AO and IO models and found no additional evidence against unit roots.

Table 3
Sequential AO trend break model

	$k = 1$			$k = \kappa(t\text{-stat})$			
Region	Break year	α	t_α	Break year	α	t_α	k
New England	1940	− 0.26	− 4.84*	1939	− 0.30	− 4.74	6
Mideast	1938	− 0.25	− 4.44	1940	− 0.57	− 4.91*	8
Great Lakes	1945	− 0.53	− 4.65	1948	− 1.32	− 5.56**	6
Plains	1947	− 0.67	− 4.86*	1947	− 0.69	− 6.11***	0
Southeast	1939	− 0.45	− 5.40**	1939	− 0.41	− 3.75	7
Southwest	1939	− 0.29	− 3.76	1940	− 0.62	− 5.30**	7
R. Mountains	1950	− 0.53	− 4.37	1951	− 0.49	− 4.57	0
Far West	1942	− 0.53	− 5.16**	1942	− 0.53	− 5.16**	1

Critical values

1 percent	5 percent	10 percent
− 5.57	− 5.08	− 4.82

***, **, and * denote statistical significance at the 1%, 5%, and 10% levels, respectively.

than found by CM. While none of the regions has a break year in 1946, all but one occur between 1938 and 1951.[11]

3. Conclusion

This paper investigates further the question posed by Carlino and Mills (1993), namely are U.S. regional incomes converging. Using time-series techniques that allow us to endogenize both the break date and the lag length, features not found in CM, we are able to significantly strengthen their results on stochastic convergence for U.S. relative regional incomes. Specifically, whereas CM are able to reject the unit root, and hence find evidence for stochastic convergence for three out of eight BEA regions, we are able to do so for seven of the eight regions and with higher degrees of significance as well.

[11]It would be interesting to see if the unit root null would be rejected by a model which, while allowing the break years to be selected endogenously, constrained them to be the same across regions. Unfortunately, while Levin and Lin (1992) consider unit roots in panel data, we are not aware of any such work which allows for trend breaks.

Besides according well with Barro and Sala-i-Martin's (1992) and CM's evidence on conditional convergence for U.S. states and regions, our results are also intuitively appealing since the neoclassical growth model makes its strongest case for convergence when considering economies that are highly similar. Hence, our results can usefully be thought of as providing a benchmark case against which the existence of convergence among other sets of similar economies may be compared.

References

Banerjee, A., R. Lumsdaine, and J. Stock, 1992, Recursive and sequential tests of the unit root and trend break hypothesis: Theory and international evidence, Journal of Business and Economic Statistics 10, 271–287.

Barro, R.J., 1992, Economic growth in a cross-section of countries, Quarterly Journal of Economics 106, 407–443.

Barro, R.J. and X. Sala-i-Martin, 1991, Convergence across states and regions, Brookings Papers on Economic Activity, 1, 107–182.

Barro, R.J. and X. Sala-i-Martin, 1992, Convergence, Journal of Political Economy 100, 223–251.

Ben-David, Dan, 1994, Income disparity among countries and the effects of freer trade, in: L.L. Pasinetti and R.M. Solow, eds., Economic growth and the structure of long run development (Macmillan, London) 45–64.

Bernard, A.B. and S.N. Durlauf, 1995, Convergence in international output, Journal of Applied Econometrics 10, 97–108.

Bernard, A.B. and S.N. Durlauf, 1996, Interpreting tests of the convergence hypothesis, Journal of Econometrics 71, 161–173.

Brown, S.J., N.E. Coulson, and R.F. Engle, 1990, Non-cointegration and econometric evaluation of models of regional shift and share, Working paper no. 3291 (NBER, Cambridge, MA).

Campbell, J.Y. and N.G. Mankiw, 1989, International evidence on the persistence of economic fluctuations, Journal of Monetary Economics 23, 319–333.

Campbell, J.Y. and P. Perron, 1991, What macroeconomists should know about unit roots, in: O.J. Blanchard and S. Fischer, eds., NBER macroeconomics annual 1991 (MIT Press, Cambridge, MA) 141–201.

Carlino, G.A. and L.O. Mills, 1993, Are U.S. regional incomes converging? A time series analysis, Journal of Monetary Economics 32, 335–346.

Christiano, L.J., 1992, Searching for a break in GNP, Journal of Business and Economic Statistics 10, 237–250.

Levin, A. and C.F. Lin, 1992, Unit root tests in panel data: Asymptotic and finite-sample properties, Discussion paper no. 92–23 (University of California, San Diego, San Diego, CA).

Lucas, R.E., Jr., 1988, On the mechanics of economic growth, Journal of Monetary Economics 22, 3–42.

MacKinnon, J.G., 1991, Critical values for cointegration tests, in: R.F. Engle and C.W.J. Granger, eds., Readings in cointegration (Oxford University Press, New York, NY) 267–276.

Mankiw, N.G., D. Romer, and D.N. Weil, 1992, A contribution to the empirics of economic growth, Quarterly Journal of Economics 107, 407–438.

Nelson, C.R. and C.I. Plosser, 1982, Trends and random walks in macroeconomic time series: Some evidence and implications, Journal of Monetary Economics 10, 139–162.

Ng, S. and P. Perron, 1995, Unit root tests in ARMA models with data dependent methods for the selection of the truncation lag, Journal of the American Statistical Association 90, 268–281.

Perron, P., 1989, The great crash, the oil price shock, and the unit root hypothesis, Econometrica 57, 1361–1401.

Perron, P., 1994, Further evidence on breaking trend functions in macroeconomic variables, Working paper (University of Montreal, Montreal, PQ).

Quah, Danny, 1990, International patterns of growth I: Persistence in cross-country disparities, Working paper (MIT, Cambridge, MA).

Romer, P.M. 1986, Increasing returns and long run growth, Journal of Political Economy 94, 1002–1037.

Solow, R.M., 1956, A contribution to the theory of economic growth, Quarterly Journal of Economics 70, 65–94.

Vogelsang, T.J. and P. Perron, 1994, Additional tests for a unit root allowing for a break in the trend function at an unknown time, Working paper no. 94-13 (Cornell University, Ithaca, NY).

Zivot, E. and D. Andrews, 1992, Further evidence on the great crash, the oil price shock, and the unit root hypothesis, Journal of Business and Economic Statistics 10, 251–270.