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REGULATORY FEDERALISM AND THE DISTRIBUTION OF AIR POLLUTANT EMISSIONS*

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ABSTRACT. Recent empirical work suggests that (i) incomes are converging through time, and (ii) income and pollution levels are linked. This paper weds these two literatures by examining the spatial and temporal distribution of pollution. After establishing that theoretical predictions about whether pollution will converge are critically linked to certain structural parameters, we explore pollution convergence using state-level data on two important pollutants—nitrogen oxides and sulfur oxides—from 1929 to 1999. We find stronger evidence of converging emission rates during the federal pollution control years (1970–1999) than during the local control years (1929–1969). These results suggest that income convergence alone may not be sufficient to induce convergence of pollutant emissions.

1. INTRODUCTION

Examining the appropriate division of functions among different levels of government has increased in popularity during the past two decades. In academic circles, Musgrave's (1959) seminal treatment of the public sector served to stimulate interest in the federalist nature of governmental structures some four decades ago. Musgrave divided the public fiscal department into three branches: allocation, distribution, and stabilization; but perhaps the area attracting the most attention of late is the allocation branch—determining the optimal institutional arrangements for the provision of public goods and services (see, e.g.,

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Oates, 1991, 2002). Within this large scope of duties lies the responsibility for setting and enforcing environmental regulations.

A hotly debated issue revolves around the merits of regulation at each level: some conjecture that in a second-best world where initial distortions are present, locally determined environmental regulations are likely to be suboptimal when jurisdictions compete with each other to attract capital (Oates and Schwab, 1988). Alternatively, others (notably Wilson, 1996) provide theoretical arguments that suggest localities may *either* "race to the top" (the NIMBY phenomenon) or "race to the bottom" when setting environmental regulations. Regulatory federalism issues are much more than academic curiosity, as concern over a race to the bottom led the U.S. government to create the environmental protection agency (EPA) in 1970 and to pass subsequent legislation establishing *national* ambient standards for air and water quality (List and Gerking, 2000).

The ongoing debate also has implications beyond policy circles: in both a positive and normative perspective the issue has considerable merits. First, equity may be important. Although a literature is developing that examines the distributional characteristics of pollution (see, e.g., Brooks and Sethi, 1997; Arora and Cason, 1999; List, 1999; Millimet and List, 2003; Strazicich and List, 2003), little has been done to link the effects of centralization (or devolution) of environmental authority to the distribution of pollutant emissions. Considerable evidence has mounted that suggests certain pollutants have threshold effects (e.g., Arrow et al., 1995), and therefore health affects from higher pollution levels are best represented by, for example, an exponential function. Accordingly, ceteris paribus, an optimal spatial pattern of pollution would favor dilution rather than agglomeration.

Second, Solow's (1956) neoclassical growth model implies that given similar economy characteristics of population growth, savings, depreciation, and technology, per capita incomes will converge across countries and regions. Given permanent differences in these characteristics, the Solow model predicts that incomes will converge "conditionally" to their own steady-state, or "compensating differential." Thus, the Solow model predicts that relatively poor regions will catch up to relatively rich regions, although permanent income differences will likely remain. Recent empirical evidence suggests that incomes are indeed converging across U.S. states and regions (e.g., Barro and Sala-i-Martin, 1991; Carlino and Mills, 1993, 1996; Evans and Karras, 1996a,b; Loewy and Papell, 1996; Quah, 1996; Webber, White, and Allen, 2005). If one links incomes and environmental outcomes via the Coase theorem (Hamilton, 1995), these findings leave open the possibility that income levels adjusted for environmental quality will also converge (see also Brock and Taylor, 2004).

The theoretical perspective on these matters is ambiguous. Even in the context of a simple Solow growth model one might find converging or diverging pollution levels. Adding complicating (but realistic) features like interjurisdictional competition for capital adds to the potentially rich set of results that may be obtained in theory. With such conflicting predictions from theory, we turn to an empirical analysis to explore whether pollutant emissions are converging across the U.S. states in the 20th century. We present evidence concerning these issues by examining 1929–1999 state-level data on emissions of sulfur dioxides and nitrogen oxides. To provide an indication of the effects of centralization on the distribution of emissions, we split the sample into two regimes—1929–1969 and 1970–1999—and use time series tests with structural breaks to test for converging emissions. The early sample period represents an era of state and local pollution control whereas the latter years correspond to an era of significant federal presence.[1]

We find that during the years of state pollution control, emissions of pollutants for about one-half of the states are not converging. Interestingly, after centralization of environmental standards, emissions are converging in about three-quarters of the states. These results are consistent with the Solow model described in Section 2, but also suggest the possibility of environmental dumping—competition between localities to attract capital investment (see, e.g., Jeppesen, List, and Folmer, 2002). According to this perspective, poorer regions may have traded environmental quality for relative income gains prior to 1970, but were restricted in their ability to do so after 1970. Regardless of interpretation, importantly, this result suggests that convergence of income may not be sufficient to induce convergence of some pollutant types. Our findings are consistent with the conjecture of Brooks and Sethi (1997) that "without [nation-wide ambient] standards, the disparities faced by certain subpopulations in the United States will not diminish appreciably in the foreseeable future." Although our data are aggregate in nature, whereas Brooks and Sethi (1997) use disaggregate data, it is interesting that their intuition is consistent with our empirical results.

The remainder of our paper is organized as follows. Section 2 sketches the main linkages between income growth and pollution when environmental regulation is implemented at the state and federal level. Section 3 describes the data. Section 4 presents the empirical methods and results. Section 5 provides concluding remarks.

2. POLLUTION PATTERNS AND CONVERGENCE

An interesting body of literature examining the determinants of spatial pollution patterns has emerged recently (Gianessi, Peskin, and Wolff, 1979;

[1]In the U.S., responsibility of regulating polluters rested almost exclusively with the states until 1970 (Portney, 1990). The federal government began to take a more active role in environmental regulation during the Nixon years when the President declared himself an environmentalist and proclaimed the onset of the "environmental decade" on January 1, 1970. President Nixon subsequently passed the most important air pollution control bill in history—the Clean Air Act Amendments of 1970—and created the EPA by executive order in 1970 to implement the Amendments. This regime change affords us the opportunity to examine pollution paths in the pre- and post-centralization eras to make inference regarding the relationship between institutional arrangements and the spatial distribution of pollutants.

Hamilton, 1995; Brooks and Sethi, 1997; Arora and Cason, 1999). Adding to this literature, we note that there are two distinct empirical observations that are germane to the issue of pollution convergence. First, as mentioned in the previous section, state and regional income levels are generally converging over time. Second, there is evidence of an inverted U-shaped relation between income and certain pollutants—the Environmental Kuznets Curve (EKC).[2] Grossman and Krueger (1995) provide some support in favor of the EKC for local stock pollutants, and List and Gallet (1999) and Millimet et al. (2003) find some evidence of such a relationship for emissions of NO_x and SO_2 across U.S. states (we should note, however, that Azomahou, Laisney, and Nguyen Van (2006) use a nonparametric panel data model to show that this nonmonotonic relation does not hold for CO_2—a global stock pollutant). The combination of income convergence and the EKC necessarily implies that theory produces ambiguous predictions about whether pollution profiles are converging or diverging.

The intuition is simple (a formal derivation is available in the Appendix). Consider a simple Solow growth model where pollution is a byproduct of consumption—but may be mitigated by abatement—and where a "mechanism" is built into the model to ensure an inverted U-shaped relation between income and pollution. Following the literature, this mechanism may be driven by political economy considerations, increasing returns to scale in abatement, or technical change. With equilibrium income growth governed by exogenous technical change in production, there must be absolute convergence of pollution in the *long term*. Pollution levels may fall to zero if incomes get sufficiently high or approach arbitrary low levels, but eventually they will converge.

The "not-so-long-term" perspective is rather different, however. Absolute pollution levels may converge or diverge—both during the approach to the balanced growth paths and along balanced growth paths. States with different income levels will converge towards their own balanced income and pollution path, but this implies that for some states (with "low" incomes) the growth rate of pollutants will be positive whereas for other states (with "high" incomes) the growth rate is negative. Pollution levels may converge or diverge, depending on income differences between states. When such income differences are small, convergence will occur; when differences are large, pollution levels may diverge.

We will briefly discuss two extensions of this basic story, highlighting that the picture becomes even less clear when adding further realistic elements to the model. First, we may consider the case where environmental regulation is the responsibility of the federal government rather than the state, as has been

[2]The causal mechanism for this EKC has been the subject of some debate. Some analysts have pointed to political economy linkages (such as based on the Coase theorem), and others have emphasized the technical relationship between consumption and abatement (e.g., Andreoni and Levinson, 2001), or changes in this relationship over time due to technical change (e.g., Stokey, 1998). Brock and Taylor (2004) present a model wherein technical change in abatement is proportional to the average abatement intensity in the economy (i.e., individual abatement efforts provide knowledge spillovers to others in the economy).

effectively the case in the U.S. since 1970.[3] As shown by Andreoni and Levinson (2001), when the federal government maximizes the sum of state utilities and accounts for spillover costs, the shape of the pollution/income path does not change. While optimal consumption (abatement) levels are lower (higher) from a federal perspective when compared to the local optimum (so that the *level* of the pollution/income path is affected); the pollution path's *shape* does not change. In the long run absolute convergence is expected; in the short and medium run convergence towards the balanced pollution path occurs, and absolute pollution levels in different states may diverge or converge.

Second, we may consider what happens when capital is a mobile production factor. In a first-best world, states will not lower taxes to attract capital—doing so would reduce local welfare and induce Tiebout migration (where residents "vote with their feet"; see Tiebout, 1956, but also Oates and Schwab, 1988, and Wellisch, 2000).[4] Real-life politics, however, takes place in a context littered with pre-existing distortions. Key assumptions of the benchmark model are likely violated in reality. Examples include lack of market power in national capital markets, absence of strategic interaction with respect to other localities, and access to a perfect range of regulatory instruments. In such a context, "optimal" local regulation might well be different from the first-best benchmark, as environmental policies should take distortions into account.

For example, competition for capital in a world in which the social return to capital exceeds the private return could give rise to sub-optimally low taxes. This might be caused by capital taxation, but there are many other potential distortions. For example, in the case of unemployment due to wage rigidities, it may be rational for governments to choose "low" tax levels to reduce unemployment costs (Wilson, 1996).[5] If tighter environmental regulation decreases the rate of return on investment, then Krugman's (1997) quip becomes important: "any country which has a pre-existing tendency to attract too little capital will have an incentive to avoid such regulations; whereas a collective, international decision to impose higher standards would not lead to capital flight, because the capital would have nowhere to go."[6]

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[3] One oft-cited disadvantage of local regulatory decision making is that spill-overs are not internalized, and therefore too much pollution is a general outcome when localities have regulatory authority. Another disadvantage of local pollution control relates to the potentially adverse consequences of interjurisdictional competition (Wilson (1996) and below).

[4] Of course, given the mobility opportunities in the U.S., individuals will migrate in response to spatial arbitrage opportunities, leading to the marginal person being indifferent to locations at any point in time.

[5] Policymakers may be tempted to resort to lax environmental policies in an effort to attract mobile capital (as opposed to, for example, more efficient capital subsidies) if environmental policy formulation is less transparent to voters ("optimal obfuscation").

[6] For an overview of reasons why competing local jurisdictions may produce distorted outcomes, refer to surveys by Wellisch (2000) and Wilson (1996, 1999). Oates (2002) concludes that the theoretical literature is inconclusive on whether races to the bottom or optimal regulation (i.e., convergence of pollution in light of the evidence above) will occur.

The theoretical insight that local policies may be distorted is consonant with the following quote, which highlights that state policymakers will trade off environmental quality to attract manufacturing firms and jobs:

> "We have traded the environment for jobs. . .where the environment became either totally or partially damaged, in some instances permanently. However, we have no regrets, no remorse."—Edwin Edwards, Governor of Louisiana, 1979.

If states in a second-best context compete for mobile capital and, in terms of environmental stringency, "race to the bottom," local economies are not on a single equilibrium income and pollution path. Specifically, when local standards or taxes are relaxed, the economy shifts to an increasingly higher pollution path. If only a subset of states is engaged in competition, then *divergence of pollution* is expected to occur. In this case, some states willingly "drift" from one pollution/income path to the next while others follow a single path over time. This provides an alternative prediction for the intertemporal behavior of pollution paths.

Writing about the potential efficiency losses associated with interjurisdictional competition, Oates (2002) concludes that "we are badly in need of empirical estimates of these distortions." The analysis below attempts to provide a first step in this direction.

3. DATA DESCRIPTION

We empirically analyze the spatial distribution of air pollution emissions by examining the time-series convergence proposition suggested in Carlino and Mills (1993). Described more fully below, we first examine the natural log of the ratio of state per capita emissions relative to average per capita emissions to test if shocks are permanent or temporary.[7] If shocks to the log ratio are temporary, then the series is "stochastically converging;" alternatively, if shocks are permanent, then divergence is evident. As noted by Carlino and Mills (1993), although stochastic convergence is necessary for convergence to occur, it is not sufficient.[8] They note that the log relative series should also be examined to see if trending towards the mean. Since this time series test for "β-convergence" does not include a measure of the initial value of the log emissions series on the right-hand side of the regression, the test is not subject to criticisms noted

[7] This type of convergence has been termed conditional convergence by Mankiw, Romer, and Weil (1992). Conditional convergence occurs if state-level emissions converge to a (state-specific) constant differential. This definition appears appropriate given that assimilative capacities, preferences, etc., differ across states.

[8] Stochastic convergence not necessarily implying convergence is further apparent from the possibility that the log ratio series may be trend-stationary along a trend that diverges from the national mean.

for the cross-section test.[9] Using each testing procedure, we examine both pre- and post-1970 emission rates.

To test for converging emissions, we analyze state per capita emissions of two criteria air pollutants: sulfur dioxide (SO_2) and nitrogen oxide (NO_x) (see also List, 1999; List and Gallet, 1999; and Millimet et al., 2003 who make use of these data). These choices are mainly driven by data considerations, as no other emissions data that we are aware of span from 1929 to 1999, but they also offer interesting pollutants to examine. The original Clean Air Act Amendments of 1970 targeted SO_2 early on due to its adverse health effects on humans and its negative impact through the deposition of acid rain. Quantified health effects associated with prolonged exposures to SO_2 include changes in pulmonary functions, respiratory ailments, and in extreme cases premature death. SO_2 can also significantly damage crops and man-made structures; total global damage estimates range in the hundreds of millions of dollars (U.S. EPA, 1997). These and other deleterious effects associated with SO_2 emissions led policymakers to take an early strong stance against emitters of SO_2 in an effort to substantially curb emissions (Gianessi et al., 1979). NO_x, on the other hand, is a precursor to ozone, can induce respiratory effects in humans, form particulate nitrates, and is a component of acid rain.

Our emissions data are available for each of the 48 contiguous states over the period 1929–1994 and come from *National Air Pollutant Emission Trends* (U.S. EPA, 1994). The EPA kindly provided additional data covering the period 1995–1999. The emission-estimating methodologies fall into two categories. From 1929–1984 emissions are calculated using a "top-down" approach, whereby state-level information on different fuel type use was entered in an algorithm to determine the various pollution emissions. For example, state level data on coal consumption was entered in an "emission factor" equation to determine the pounds of SO_2 produced for each ton of coal consumed. Emission factors were adjusted by the EPA for differences in age and type of utility, among other factors, when possible. Prior to the 1970s, there were fewer differences in emissions due to age of utility and control factors. While the primary fuel consumption measure was at the state-level, in some cases such data was not available. In this case, national fuel consumption data was employed to estimate state-level emissions as a fraction of the national.[10] From 1985 to 1999, emissions are estimated using a "bottom-up" methodology whereby emissions are derived at the plant or county level and aggregated to the state level. Note

[9]For examples of papers that criticize the cross-section test of β-convergence see Friedman, 1992; Quah, 1993; Evans and Karras, 1996a.

[10]For a detailed discussion of the methods used to estimate our emissions data please see the document "National Air Pollutant Emission Trends, Procedures Document, 1900–1996" (U.S. Environmental Protection Agency, 1998). Various procedure reports are available on the web at "http://www.epa.gov/ttn/chief/trends/index.html." It should be noted that the emissions data employed in our empirical tests are primarily estimated from actual fuel consumption and emissions, and not from state or national aggregate income, industrial production, or GDP. We thank Roy Huntley and the staff of the EPA for helpful discussions regarding the emissions data.

TABLE 1: Aggregate U.S. SO_2 and NO_x Emissions (in Thousands of Short Tons)

Year	SO_2 emissions	NO_x emissions
1970	31,161	20,625
1971	29,686	21,227
1972	30,389	22,381
1973	31,756	23,118
1974	30,035	22,317
1975	28,011	21,889
1976	28,426	23,258
1977	28,605	23,986
1978	26,899	24,253
1979	26,953	23,812
1980	25,905	23,281
1981	24,527	23,070
1982	23,140	22,496
1983	22,541	22,364
1984	23,470	23,172
1985	23,230	22,859
1986	22,442	22,348
1987	22,204	22,403
1988	22,647	23,618
1989	22,785	23,222
1990	22,433	23,038
1991	22,068	22,672
1992	21,836	22,847
1993	21,517	23,276
1994	21,118	23,615

that utilizing data relative to the mean, and allowing for structural breaks, mitigates problems associated with combining two potentially heterogeneous data sets.

To indicate the degree of progress that has been made to curb emissions in the last three decades, we present Table 1, which provides a breakdown of the *national* emissions estimates from 1970 to 1994. The data presented show that SO_2 emissions were significantly declining after federal intervention: between 1970 and 1994 raw SO_2 emissions decreased from 31,161 thousand short tons to 21,118 thousand short tons, a reduction of nearly 33 percent. Alternatively, while NO_x emissions increased by over 180 percent from 1940 to 1970, emissions since 1970 have leveled off at about 23 million short tons since 1973. The result for SO_2 is consistent with our prediction that centralization lowers aggregate emissions; while the data show that centralization did not decrease the level of NO_x, it did significantly lower its growth rate. Yet the limited progress that has been made to reduce NO_x emissions highlights why ozone has attracted the most regulatory attention of late. While these data patterns are interesting in their own right, they provide little guidance about the temporal distribution of pollutants. To examine whether emissions are converging, further testing using state-level data is necessary.

4. EMPIRICAL METHODOLOGY AND RESULTS

Stochastic Convergence

The null hypothesis that state emission levels are diverging is first examined by testing for a unit root in the logarithm of pollution per capita in state i relative to the average of all states. Stochastic convergence implies that shocks to state pollution relative to the average will be temporary, implying that the log-relative series is a stationary process. Following a shock to the system, per capita pollution in state i can move further above or below the national average only temporarily. Alternatively, under the unit root null hypothesis, relative emissions are nonstationary, suggesting that shocks have permanent effects. Following a shock to per capita emissions in state i, under the null hypothesis, there is no tendency to return toward the average, implying that emissions will diverge. Since a state-specific constant term is included in the test, stochastic convergence is consistent with conditional convergence.

To illustrate, we let x_{it} denote the natural logarithm of the ratio of state i per capita pollution relative to the national average in year t. A unit root implies that x_{it} is a nonstationary series, which can be described as follows

$$(1) \qquad x_{it} = \mu_i + \beta_i x_{it-1} + \varepsilon_{it},\ \varepsilon_{it} \sim (0, \sigma^2) \quad \text{and} \quad E(\varepsilon_{it}\varepsilon_{it-j}) = 0,\ j \neq 0,$$

where μ_i is a state-specific constant term or "drift," $\beta_i = 1$, and ε_{it} is a white noise error term that is independent and identically distributed. A nonstationary x_{it} implies that any new shock (ε_{it}) to relative pollution in state i will cause a permanent change with no tendency to converge back to a stable state-specific compensating differential, implying that per capita emission levels will diverge.

Perron (1989) demonstrates that failure to allow for an existing structural break leads to a bias against rejecting a false unit root null hypothesis. Perron (1989) proposes to allow for one known, or "exogenous," structural break in the augmented Dickey-Fuller (ADF, hereafter) unit root test. Following Perron (1989), Zivot and Andrews (1992) (ZA hereafter), and others, propose ADF-type tests to determine the break point "endogenously" from the data. The ZA test selects the break point where the t-statistic that tests the unit root null is minimized (i.e., the most negative) and, therefore, least favorable to the unit root hypothesis.

A potential problem common to the ZA and other ADF-type endogenous break unit root tests is that they derive their critical values while assuming no structural break(s) under the null. Nunes, Newbold, and Kuan (1997) and Lee and Strazicich (2001) show that this assumption leads to size distortions in the presence of a unit root with break. When utilizing ADF-type endogenous break tests, researchers might incorrectly conclude that a time series is trend-stationary when in fact the series is nonstationary with break(s). As such, a "spurious rejection" may result. To provide a remedy, we utilize the minimum Lagrange multiplier (LM) unit root tests of Lee and Strazicich (2003, 2004). These tests endogenously determine one or two structural breaks in level and trend.

Implementation of the two-break minimum LM unit root test can be described as follows. According to the LM (score) principle, a unit root test statistic can be obtained from the following regression

$$\Delta x_t = \delta' \Delta Z_t + \phi \tilde{S}_{t-1} + \Sigma \gamma_i \Delta \tilde{S}_{t-i} + \varepsilon_t, \tag{2}$$

where $\tilde{S}_t$ is a de-trended series such that $\tilde{S}_t = x_t - \tilde{\Psi}_x - Z_t\tilde{\delta}, t = 2, \ldots, T$; $\tilde{\delta}$ is a vector of coefficients in the regression of Δx_t on ΔZ_t and $\tilde{\Psi}_x = x_1 - Z_1\tilde{\delta}$, where Z_t is defined below; x_1 and Z_1 are the first observations of x_t and Z_t, respectively; Δ is the difference operator; and ε_t is the contemporaneous error term and is assumed independent and identically distributed with zero mean and finite variance. To correct for serial correlation, k number of lagged augmented terms $\Delta \tilde{S}_{t-1}, i = 1, \ldots, k$, are included as necessary.[11] Z_t is a vector of exogenous variables defined by the data-generating process. With two structural breaks in intercept and trend, Z_t can be described as $[1, t, D_{1t}, D_{2t}, DT_{1t}{}^*, DT_{2t}{}^*]'$, where $D_{jt} = 1$ for $t \geq T_{Bj} + 1, j = 1, 2$, and zero otherwise, $DT_{jt}{}^* = t - T_{Bj}$ for $t \geq T_{Bj} + 1, j = 1, 2$, and zero otherwise, and T_{Bj} denotes the time period of the breaks. Note that the test regression (2) utilizes ΔZ_t instead of Z_t so that ΔZ_t becomes $[1, B_{1t}, B_{2t}, D_{1t}, D_{2t}]'$, where $B_{jt} = \Delta D_{jt}$ and $D_{jt} = \Delta DT_{jt}{}^*, j = 1, 2$. Under the unit root null hypothesis, $\phi = 0$ in equation (2), and the test statistic can be defined as

$$\tilde{\tau} = t - \text{statistic for the null hypothesis } \phi = 0. \tag{3}$$

To determine endogenously the location of the two breaks ($\lambda_j = T_{Bj}/T, j = 1, 2$), the minimum LM unit root test finds the combination of breaks where the t-statistic testing the unit root null hypothesis is minimized. Since the critical values for the minimum LM test with level and trend breaks depend (somewhat) on the location of the breaks, we use critical values that correspond to λ_j.[12]

[11]To determine the optimal number of k lagged augmentation terms that correct for serial correlation in equation (2) we follow the "general to specific" procedure suggested by Perron (1989). At each combination of two breaks in the time interval [.1T, .9T] (to eliminate end points), where T is the sample size, we begin with a maximum number of lagged first-differenced terms $k = 8$ and examine the last term to see if it is significantly different from zero at the 10-percent level in an asymptotic normal distribution (critical value is 1.645). If insignificant, the maximum lagged term is dropped and the model is reestimated with $k = 7$ terms and so on, until either the maximum term is found or $k = 0$, at which point the procedure stops. This technique has been shown to perform well as compared to other data dependent procedures to select the number of lagged augmented terms (e.g., Ng and Perron, 1995).

[12]Although the critical values for the minimum LM test with intercept and trend breaks depend (somewhat) on the location of the break(s), the test remains free of size distortions and spurious rejections in the presence of a unit root with break(s).

After identifying two breaks in level and trend for each state, we examine the significance of each estimated break coefficient at the 10-percent level in an asymptotic normal distribution (i.e., critical value is 1.645). If two breaks are not significant, we reestimate our data using the one-break minimum LM unit root test of Lee and Strazicich (2004). If no break is significant at the 10-percent level in the one-break test, we utilize the conventional (no-break) ADF test.[13]

Time Series Test for β-*Convergence*

As previously noted, the conventional cross-section test for β-convergence has been criticized in the literature. In this paper, we utilize a time series test for β-convergence that was suggested in Carlino and Mills (1993). If per capita incomes are converging, then regression of the log relative income series on an intercept and trend should find opposite signs on their estimated coefficients. Intuitively, if a state has initial pollution emissions above the national mean then the intercept term of the log-relative series would be positive and the rate of growth of emissions should be negative in order for convergence to occur. Given that we identify structural breaks in level and trend for nearly all states, we wish to utilize this information in our time series tests for β-convergence.

Our testing methodology can be described as follows. After performing unit root tests on the log of relative emissions as described above, we identify those states that reject the unit root null hypothesis (at the 10 percent level of significance). These time-series can be described as (trend) stationary around (in nearly all cases) one or two structural breaks in level and trend. Then, OLS regressions are performed on the log-relative emission series (x_t) for each of these states as follows

$$x_t = \mu_1 D1_t + \mu_2 D2_t + \mu_3 D3_t + \gamma_1 \mathrm{TIME1}_t + \gamma_2 \mathrm{TIME2}_t + \gamma_3 \mathrm{TIME3}_t + \varepsilon_t, \tag{15}$$

where $D1_t$, $D2_t$, and $D3_t$, are intercept dummy variables and $TIME1_t$, $TIME2_t$, and $TIME3_t$ are linear time trends. $D1_t = 1$ if $t \leq T_{B1}$, and zero otherwise; $D2_t = 1$ if $T_{B1} < t \leq T_{B2}$, and zero otherwise; and $D3_t = 1$ if $t > T_{B2}$, and zero otherwise. $TIME1_t = t$ if $t \leq T_{B1}$, and zero otherwise; $TIME2_t = t - T_{B1}$ if $T_{B1} < t \leq T_{B2}$, and zero otherwise; and $TIME3_t = t - T_{B2}$ if $t > T_{B2}$, and zero otherwise. Thus, μ_1, μ_2, and μ_3 denote whether the level of emissions in state i are above (+) or below (–) the national average emissions prior to the first break, following the first break, and following the second break, respectively. In a similar manner, γ_1, γ_2, and γ_3 denote the growth rate of emissions relative to the mean over these same time periods, respectively.

Prior to 1970 we are interested in the time period following the final break to determine if emissions are converging, so we compare the signs of the

[13]Gauss codes for the one- and two-break minimum LM unit root test are available online at http://www.cba.ua.edu/~jlee/gauss.

estimated coefficients μ_3 and γ_3. If the level of emissions in state i is converging to the mean, then μ_3 and γ_3 should be of opposite sign.[14] For states with only one significant break, equation (4) would be estimated without the $D3_t$ and $TIME3_t$ terms and convergence would imply opposite signs on μ_2 and γ_2 (i.e., opposite signs on level and trend). For states with no significant break, equation (4) would be estimated with only $D1_t$ and $TIME1_t$ and convergence would imply opposite signs on μ_1 and γ_1. For the later time period of 1970–1999 we initially focus on the period immediately after 1970 (i.e., μ_1 and γ_1) to examine the short-run impact of federal controls. Since changing trends in spatial pollution patterns may take time, we will also examine the long-run effect on convergence by comparing signs on the intercept and trend following the final identified structural break (i.e., μ_3 and γ_3). In addition, if the intercept term is significantly different from zero, but the trend is not, this would indicate evidence that emissions have conditionally converged to a state-specific constant differential.[15] If neither the intercept nor the trend were significantly different from zero, this would indicate convergence is unconditional, or "absolute." In contrast, if the intercept and trend is the same sign then this would suggest that emissions are diverging. If the intercept term is not significantly different from zero, but the trend is, this would also suggest that emissions are diverging.[16]

Empirical Results

Prior to testing for convergence, it is interesting to examine our per capita emissions data to determine if the period around 1970 can be identified as a time of structural change. To perform this task, we estimate the minimum LM unit root test for each pollutant type using the raw data for the entire sample period of 1929–1999. The results are displayed in Table 2. The results indicate that the period surrounding 1970 does indeed contain a significant number of structural breaks in emissions. For each pollutant type, our results identify significant structural breaks in 21 (44 percent) of 48 states during the period around 1970 (1968–1976). We will now examine the time periods before and after 1970.

Lack of Convergence in Most States under Local Control

Estimation results are presented in Tables 3–6. We first examine the test results for stochastic convergence in Tables 3 and 4. Regardless of the pollutant type and time period examined, the null of a unit root is rejected in nearly all cases (i.e., for SO_2 85 percent of states reject the unit root null before 1970 and

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[14]See Tomljanovich and Vogelsang (2002) for a similar application of the time-series notion of β-convergence applied to per capita incomes of the U.S. regions.

[15]Tomljanovich and Vogelsang (2002) do not consider this possibility.

[16]Ibid.

TABLE 2: Minimum LM Unit Root Tests of Per Capita Emissions

	SO_2		NO_x	
State	1929–1999	Breaks	1929–1999	Breaks
Alabama	−5.13	1953, 1970	−3.51	1962
Arizona	−6.27**	1943, 1976	−7.38***	1964, 1981
Arkansas	−5.02	1956, 1979	−2.97	1943
California	−6.12**	1942, 1950	−4.31	1949, 1976
Colorado	−6.51***	1946, 1961	−5.53*	1942, 1974
Connecticut	−5.37*	1956, 1971	−4.71**	1965
Delaware	−5.13	1956, 1981	−4.74	1977, 1992
Florida	−4.55	1969, 1980	−8.11***	1951, 1969
Georgia	−5.11	1968, 1984	−4.09	1945, 1980
Idaho	−5.96**	1944, 1972	−7.91***	1939, 1975
Illinois	−7.51***	1951, 1956	−4.27	1939, 1975
Indiana	−5.55*	1951, 1977	−2.99	1977
Iowa	−6.74***	1941, 1952	−5.12	1949, 1978
Kansas	−5.08	1956, 1973	−4.99	1974, 1991
Kentucky	−5.97**	1953, 1970	−5.66*	1953, 1970
Louisiana	−7.67***	1974, 1978	−5.71*	1940, 1972
Maine	−4.82	1946, 1961	−5.13	1946, 1965
Maryland	−4.78	1939, 1950	−5.00	1961, 1978
Massachusetts	−5.30*	1969, 1977	−5.18	1959, 1980
Michigan	−5.81**	1963, 1979	−4.71	1962, 1976
Minnesota	−6.69***	1946, 1966	−12.79***	1940, 1950
Mississippi	−4.99	1947, 1969	−4.33	1948, 1979
Missouri	−4.07	1971	−3.87	1973
Montana	−6.40***	1977	−5.05**	1987
Nebraska	−6.26***	1940, 1952	−5.44*	1945, 1981
Nevada	−6.75***	1961, 1979	−8.28***	1945, 1962
New Hampshire	−4.18	1946	−5.81**	1946, 1966
New Jersey	−5.89**	1939, 1950	−2.78	
New Mexico	−6.04**	1964, 1983	−5.58*	1975, 1979
New York	−5.23	1965, 1977	−5.60*	1969, 1978
North Carolina	−4.31*	1965	−4.99	1966, 1984
North Dakota	−5.86**	1974, 1992	−7.25***	1967, 1974
Ohio	−6.97***	1977	−5.25	1949, 1976
Oklahoma	−4.57	1954, 1976	−4.44*	1968
Oregon	−5.71**	1939, 1955	−7.22***	1943, 1952
Pennsylvania	−5.65*	1966, 1978	−1.05	
Rhode Island	−7.44***	1952, 1971	−6.10**	1951, 1972
South Carolina	−4.42	1956, 1977	−5.34*	1942, 1968
South Dakota	−4.47	1964, 1980	−4.84***	
Tennessee	−5.77**	1957, 1970	−4.58	1955, 1980
Texas	−4.44	1948, 1974	−5.84**	1940, 1969
Utah	−6.26**	1950, 1971	−5.33*	1948, 1970
Vermont	−5.34*	1940, 1951	−5.09***	1947
Virginia	−4.70	1941, 1952	−4.47*	1965
Washington	−4.28	1975, 1980	−4.68**	1973
West Virginia	−4.35*	1969	−5.63*	1965, 1973
Wisconsin	−6.86***	1946, 1967	−3.57	1964
Wyoming	−6.87***	1970, 1988	−7.07***	1966, 1980

Notes: The dependent variable is the level of per capita emissions. Test statistic is the t-statistic that tests the null hypothesis of a unit root. All unit root tests include an intercept and trend. Breaks denote the break years that were identified by the one- or two-break minimum LM unit root test. A blank space denotes no breaks were significant at the 10 percent level. In the case of no significant breaks, the results shown are those obtained using the conventional (no break) ADF test. ***,**, and * denote significance at the 1 percent, 5 percent, and 10 percent levels, respectively. Critical values for the one- and two-break minimum LM test come from Lee and Strazicich (2003, 2004). Critical values for the (no-break) ADF test are from MacKinnon (1991).

TABLE 3: Minimum LM Unit Root Tests for Stochastic Convergence of SO_2

State	1929–1969	Breaks	1970–1999	Breaks
Alabama	−9.76***	1950, 1957	−6.41**	1981, 1985
Arizona	−8.89***	1941, 1950	−7.17***	1987, 1991
Arkansas	−5.79**	1948, 1957	−5.06***	1987
California	−5.13***	1949	−7.18***	1984, 1994
Colorado	−5.06	1950, 1960	−7.79***	1983, 1993
Connecticut	−8.76***	1943, 1958	−5.85**	1990, 1993
Delaware	−7.93***	1945, 1957	−9.68***	1980, 1993
Florida	−9.77***	1940, 1955	−5.71**	1981, 1989
Georgia	−5.87**	1950, 1965	−6.47***	1987, 1996
Idaho	−6.94***	1947, 1960	−7.40***	1982, 1986
Illinois	−5.66*	1941, 1952	−6.07**	1988, 1996
Indiana	−7.38***	1941, 1948	−5.45*	1983, 1987
Iowa	−6.72***	1951, 1960	−8.86***	1981, 1986
Kansas	−7.97***	1951, 1955	−5.98**	1982, 1993
Kentucky	−6.99***	1950, 1958	−5.39*	1986, 1991
Louisiana	−5.92***	1957	−10.02***	1984, 1988
Maine	−5.39*	1941, 1949	−12.34***	1981, 1993
Maryland	−3.92	1939, 1953	−6.48***	1985, 1994
Massachusetts	−9.88***	1944, 1956	−7.96***	1980, 1990
Michigan	−5.96**	1944, 1964	−6.49***	1980, 1994
Minnesota	−4.62**	1953	−6.87***	1983, 1993
Mississippi	−5.57*	1946, 1955	−6.74***	1980, 1986
Missouri	−4.63	1947, 1961	−11.66***	1981, 1991
Montana	−9.34***	1942, 1955	−11.87***	1981, 1995
Nebraska	−6.93***	1951, 1955	−14.61***	1980, 1993
Nevada	−5.64*	1944, 1950	−4.83	1980, 1986
New Hampshire	−7.53***	1947, 1957	−10.32***	1982, 1993
New Jersey	−4.16	1948	−8.20***	1988, 1993
New Mexico	−6.71***	1951, 1960	−4.85	1982, 1987
New York	−5.98**	1951, 1960	−5.52***	1984
North Carolina	−6.72***	1943, 1956	−7.64***	1985, 1994
North Dakota	−6.72***	1945, 1949	−10.89***	1980, 1993
Ohio	−6.52***	1951, 1959	−7.21***	1988, 1994
Oklahoma	−5.86**	1956, 1960	−6.46***	1981, 1993
Oregon	−5.32*	1956, 1965	−5.38*	1983, 1989
Pennsylvania	−5.35*	1949, 1962	−5.98***	1988, 1994
Rhode Island	−4.58	1942, 1960	−9.15***	1982, 1993
South Carolina	−4.62	1939, 1952	−4.48	1980, 1991
South Dakota	−8.18***	1943, 1964	−8.15***	1986, 1992
Tennessee	−6.16**	1947, 1958	−7.72***	1987, 1992
Texas	−6.28***	1948	−5.48*	1983, 1992
Utah	−8.29***	1946, 1956	−6.07**	1982, 1996
Vermont	−10.37***	1947, 1954	−22.70***	1985, 1993
Virginia	−4.78	1943, 1959	−8.20***	1981, 1993
Washington	−5.52**	1944, 1952	−11.56***	1984, 1990
West Virginia	−7.37***	1947, 1955	−4.72	1983, 1994
Wisconsin	−7.24***	1947, 1956	−7.60***	1990, 1993
Wyoming	−5.35***	1962	−6.42***	1981, 1993

Notes: The dependent variable is the natural logarithm of the ratio of per capita SO_2 emissions for state i in year t divided by the average per capita SO_2 emissions of 48 states in year t. Test statistic is the t-statistic that tests the null hypothesis of a unit root. Rejection of the null hypothesis supports stochastic convergence. All unit root tests include an intercept and trend. Breaks denote the break years that were identified by the one- or two-break minimum LM unit root test. A blank space denotes no breaks were significant at the 10 percent level. In the case of no significant breaks, the results shown are those obtained using the conventional (no break) ADF test. ***,**, and * denote significance at the 1 percent, 5 percent, and 10 percent levels, respectively. Critical values for the one- and two-break minimum LM test come from Lee and Strazicich (2003, 2004). Critical values for the (no-break) ADF test are from MacKinnon (1991).

TABLE 4: Minimum LM Unit Root Tests for Stochastic Convergence of NO_x

State	1929–1969	Breaks	1970–1999	Breaks
Alabama	−6.71***	1943, 1955	−3.82	1990
Arizona	−10.22***	1941, 1951	−11.58***	1984, 1991
Arkansas	−7.52***	1941, 1953	−9.88***	1991
California	−8.26***	1958	−6.01**	1985, 1994
Colorado	−3.87	1940	−5.22	1981, 1994
Connecticut	−6.70***	1941, 1947	−9.70***	1984, 1987
Delaware	−7.11***	1943, 1955	−15.40***	1980, 1993
Florida	−8.07***	1945, 1952	−6.31**	1982, 1989
Georgia	−5.50**	1941, 1950	−6.84***	1983, 1995
Idaho	−5.95**	1941, 1945	−6.43**	1982, 1987
Illinois	−5.77**	1948, 1955	−7.60***	1981, 1991
Indiana	−5.86**	1940, 1946	−7.06***	1985
Iowa	−6.18**	1947, 1950	−7.80***	1980, 1993
Kansas	−5.80**	1941, 1956	−6.09**	1986, 1991
Kentucky	−6.14***	1952	−10.01***	1991, 1996
Louisiana	−7.89***	1942, 1957	−11.55***	1982, 1993
Maine	−4.89	1941, 1955	−5.64*	1984, 1987
Maryland	−4.08	1946	−5.97**	1984, 1993
Massachusetts	−4.80***		−5.12	1983, 1987
Michigan	−5.01***		−7.92***	1983, 1989
Minnesota	−14.59***	1940, 1949	−9.06***	1982, 1993
Mississippi	−5.25	1941, 1951	−4.73**	1983
Missouri	−4.57	1948, 1963	−9.68***	1980, 1993
Montana	−7.24***	1941, 1954	−13.63***	1987, 1996
Nebraska	−4.89**	1949	−9.93***	1988, 1993
Nevada	−7.82***	1945, 1959	−6.68***	1984, 1994
New Hampshire	−5.23***	1945	−5.92**	1980, 1984
New Jersey	−4.89**	1947	−11.27***	1993, 1996
New Mexico	−5.93**	1946, 1956	−6.97***	1980, 1990
New York	−8.60***	1947, 1957	−6.39**	1982, 1995
North Carolina	−4.93	1947, 1955	−6.66***	1986, 1995
North Dakota	−2.93		−8.74***	1982, 1993
Ohio	−5.73***	1948	−6.02**	1981, 1993
Oklahoma	−2.85		−7.17***	1987, 1994
Oregon	−5.00**	1949	−4.44	1980, 1993
Pennsylvania	−6.90***	1945, 1957	−6.17**	1982, 1989
Rhode Island	−6.07**	1947, 1958	−5.68***	1981
South Carolina	−9.31***	1939, 1947	−6.79***	1986, 1993
South Dakota	−8.06***	1940, 1961	−7.33***	1988, 1993
Tennessee	−6.51***	1944, 1952	−5.17	1981, 1993
Texas	−4.38***		−11.76***	1982, 1993
Utah	−8.90***	1946, 1957	−7.43***	1982, 1994
Vermont	−9.86***	1947, 1959	−10.95***	1982, 1993
Virginia	−6.54***	1944, 1949	−10.42***	1981, 1993
Washington	−7.69***	1947, 1954	−5.33*	1984, 1993
West Virginia	−6.63***	1942	−5.52*	1989, 1995
Wisconsin	−7.48***	1942, 1950	−8.64***	1986, 1990
Wyoming	−6.38**	1941, 1958	−4.78	1980, 1994

Notes: The dependent variable is the natural logarithm of the ratio of per capita NO_x emissions for state i in year t divided by the average per capita NO_x emissions of 48 states in year t. Test statistic is the t-statistic that tests the null hypothesis of a unit root. Rejection of the null hypothesis supports stochastic convergence. All unit root tests include an intercept and trend. Breaks denote the break years that were identified by the one- or two-break minimum LM unit root test. A blank space denotes no breaks were significant at the 10 percent level. In the case of no significant breaks, the results shown are those obtained using the conventional (no break) ADF test. ***,**, and * denote significance at the 1 percent, 5 percent, and 10 percent levels, respectively. Critical values for the one- and two-break minimum LM test come from Lee and Strazicich (2003, 2004). Critical values for the (no-break) ADF test are from MacKinnon (1991).

TABLE 5: Summary of Time-Series Tests for Convergence of SO_2 Emissions

State	Before 1970	After 1970 short-run long-run	State	Before 1970	After 1970 short-run long-run
Alabama	d	C, D	Nebraska	c	C, c
Arizona	C	C, d	Nevada	C	U
Arkansas	D	C, C	New Hampshire	C	c, d
California	D	D, c	New Jersey	U	D, c
Colorado	U	C, c	New Mexico	D	U
Connecticut	D	c, c	New York	D	C, C
Delaware	C	A, C	North Carolina	C	C, d
Florida	c	c, C	North Dakota	D	C, D
Georgia	c	C, A	Ohio	c	D, c
Idaho	C	c, c	Oklahoma	D	C, c
Illinois	C	C, c	Oregon	c	c, C
Indiana	D	D, C	Pennsylvania	D	D, c
Iowa	C	C, d	Rhode Island	U	c, c
Kansas	D	C, C	South Carolina	U	U
Kentucky	D	C, C	South Dakota	C	c, C
Louisiana	D	C, A	Tennessee	c	c, C
Maine	C	c, c	Texas	D	C, D
Maryland	U	C, A	Utah	C	C, d
Massachusetts	c	C, c	Vermont	c	c, C
Michigan	c	C, c	Virginia	U	c, A
Minnesota	C	D, C	Washington	D	C, c
Mississippi	c	C, C	West Virginia	D	U
Missouri	U	D, C	Wisconsin	D	C, C
Montana	C	C, A	Wyoming	D	D, D

Notes: To be considered as converging, the time-series of log-relative SO_2 emissions for state i must satisfy the following criteria. First, the log-relative time-series must reject the null hypothesis of a unit root, implying that emissions are not diverging. Next, for those states that reject the unit root null hypothesis (at the 10 percent level of significance), an OLS regression on the intercept(s) and trend(s) was performed that includes structural breaks identified with the minimum LM unit root test. Finally, the coefficients on the intercept and trend in the time period before and after 1969, respectively, must indicate that emissions are converging or have already converged. The short run describes the initial intercept and trend after 1969, while the long run describes the final intercept and trend. The notation is described as follows:

"C" denotes point estimates consistent with convergence that are each statistically significant at the 10 percent level (i.e., coefficient of intercept and trend are of opposite signs).

"c" denotes point estimates that suggest conditional convergence has occurred (i.e., the intercept is statistically significant at the 10 percent level and the trend is not).

"A" denotes point estimates very small in magnitude and statistically insignificant, which combined with being stationary, suggests that absolute convergence has occurred.

"D" denotes point estimates consistent with divergence that are each statistically significant at the 10 percent level (i.e., coefficient of intercept and trend have the same sign).

"d" denotes point estimates consistent with divergence, wherein the trend is statistically significant at the 10 percent level and the intercept is not.

"U" denotes unit root, which implies that the null hypothesis of divergence cannot be rejected.

Therefore, C, c, and A denote evidence of convergence and D, d, and U imply that divergence cannot be rejected.

TABLE 6: Summary of Time-Series Tests for Convergence of NO_x Emissions

State	Before 1970	After 1970 short-run long-run	State	Before 1970	After 1970 short-run long-run
Alabama	d	U	Nebraska	C	d, D
Arizona	d	d, A	Nevada	A	c, A
Arkansas	C	C, C	New Hampshire	c	D, C
California	c	D, c	New Jersey	d	D, c
Colorado	U	U	New Mexico	D	C, c
Connecticut	c	D, c	New York	c	c, c
Delaware	A	A, C	North Carolina	U	d, C
Florida	d	c, c	North Dakota	U	D, c
Georgia	c	C, A	Ohio	c	c, c
Idaho	C	C, A	Oklahoma	U	c, C
Illinois	C	d, C	Oregon	d	U
Indiana	D	C, C	Pennsylvania	A	A, c
Iowa	A	A, D	Rhode Island	D	C, C
Kansas	c	D, c	South Carolina	C	C, d
Kentucky	d	D, c	South Dakota	C	A, d
Louisiana	c	C, c	Tennessee	c	U
Maine	U	c, C	Texas	D	C, C
Maryland	U	c, c	Utah	C	C, c
Massachusetts	D	U	Vermont	c	c, C
Michigan	C	D, c	Virginia	C	D, c
Minnesota	c	C, d	Washington	c	C, c
Mississippi	U	C, C	West Virginia	D	D, c
Missouri	U	d, c	Wisconsin	C	C, A
Montana	C	D, D	Wyoming	D	U

Notes: To be considered as converging, the time-series of log-relative NO_x emissions for state i must satisfy the following criteria. First, the log-relative time-series must reject the null hypothesis of a unit root, implying that emissions are not diverging. Next, for those states that reject the unit root null hypothesis (at the 10 percent level of significance), an OLS regression on the intercept(s) and trend(s) was performed that includes structural breaks identified with the minimum LM unit root test. Finally, the coefficients on the intercept and trend in the time period before and after 1969, respectively, must indicate that emissions are converging or have already converged. The short run describes the initial intercept and trend after 1969, while the long run describes the final intercept and trend. The notation is described as follows:

"C" denotes point estimates consistent with convergence that are each statistically significant at the 10 percent level (i.e., coefficient of intercept and trend are of opposite signs).

"c" denotes point estimates that suggest conditional convergence has occurred (i.e., the intercept is statistically significant at the 10 percent level and the trend is not).

"A" denotes point estimates very small in magnitude and statistically insignificant, which combined with being stationary, suggests that absolute convergence has occurred.

"D" denotes point estimates consistent with divergence that are each statistically significant at the 10 percent level (i.e., coefficient of intercept and trend have the same sign).

"d" denotes point estimates consistent with divergence, wherein the trend is statistically significant at the 10 percent level and the intercept is not.

"U" denotes unit root, which implies that the null hypothesis of divergence cannot be rejected.

Therefore, C, c, and A denote evidence of convergence and D, d, and U imply that divergence cannot be rejected.

92 percent after; for NO_x 83 percent of states reject the unit root null before 1970 and 88 percent after).[17] These findings clearly reject the null that differences from the mean in per capita emission rates are nonstationary and, thus, provide support for stochastic convergence. As noted above, however, stochastic convergence is a necessary but not sufficient condition for convergence. To further our investigation, we next examine the results of testing the time-series notion of β-convergence in Tables 5 and 6. These results are consistent for both types of pollutants. Prior to 1970 and the period of federal pollution control, the results for SO_2 indicate that for about one-half of the states (52 percent) pollution emission rates were diverging. For NO_x, the results indicate that emissions are diverging for nearly half of the states (44 percent).

These results make sense in light of our discussion of interjurisdictional competition and the possibility of a regulatory race to the bottom. Industrial processes have historically been large emitters of SO_2. Within this group of emitters are firms in the chemicals sector, primary metals industry, paper and allied sector, and many other types of manufacturers, all of which provide a tax base (jobs, etc.) annually to state and local economies. As discussed above, in their pursuit of firms and jobs, some policymakers may have attempted to attract polluting industry, inducing a divergence of emissions (see also Break, 1967, and Cumberland, 1980, who develop this "race to the bottom" argument more fully). Combining our empirical estimates with the fact that over this time period poorer states witnessed a much larger increase in their SO_2 emission levels than wealthier states, we conclude that there is evidence which suggests some states have drifted away from one income/pollution path to another over time. This suggests the possibility that some states were setting increasingly lenient environmental policies to attract mobile capital (since in spite of converging incomes, emissions convergence is not predominate in the data).

Convergence Followed the Environmental Movement

Examining data from 1970 to 1999, a time span when the federal government played a much stronger role in air pollution regulation via the Clean Air Act Amendments of 1970, we find different results. During this later period of more federal pollution control, we see stronger evidence that emissions are converging. The post-1970 results are displayed in Tables 5 and 6. We first examine the (short-run) impact on convergence in the period immediately following the imposition of federal controls. For SO_2, in contrast to the earlier time period, pollution emissions are now clearly converging for most states (75 percent). For

[17]While using panel data can increase power in unit root tests, given the strong rejection of the unit root hypothesis for nearly all states, the use of panel tests would not provide any additional information as compared to our univariate tests. For example, using the panel unit root test of Im, Pesaran, and Shin (2003) a rejection of the null indicates only that at least one state is stochastically converging.

NO_x, the evidence is essentially unchanged as compared to the earlier period, with convergence occurring in about one-half (54 percent) of the states. Given that adjustment of pollution patterns to federal controls may take time, we next examine the long-run evidence of convergence after 1970. In the long run, the evidence for convergence of SO_2 emissions is essentially unchanged (73 percent of states are found to have converged), but the evidence for convergence of NO_x gets considerably stronger: 75 percent of states are now found to be converging (up from 54 percent).

These empirical findings suggest that an added silver lining followed the centralization of environmental policies in the U.S.: since many types of pollutants are believed to have threshold effects (e.g., a minimum level that can be emitted before serious harm), pollutants having an equivalent conditional mean across space may decrease the chance of crossing important thresholds. In addition, from a social welfare perspective, this finding is of note because policymakers may be interested in a more equal distribution of public "bads," such as pollution.[18]

5. CONCLUSIONS

Empirical studies examining the temporal path of relative incomes have recently proliferated, with the bulk of evidence suggesting that the conditional convergence prediction of Solow's (1956) neoclassical growth model adequately describes the growth process. At the same time, a large literature has developed that finds a link between income and pollution levels through time (e.g., Grossman and Krueger, 1995). While these findings are important in isolation, combining their key features yields potentially significant new insights. In this study we link these two literatures by examining the spatial and temporal distribution of pollution. Since key aspects of any public policy are efficiency and distributional consequences, our results may also have policy relevance.

Using state-level data on sulfur dioxide and nitrogen oxide emissions from 1929 to 1999, we find that during the years of state pollution control, for about one-half of the states, emissions did not converge. Interestingly, lower state emission levels and stronger evidence of emissions convergence occurred in the era of federal controls. These findings are consonant with the notion that without a nationwide pollution control standard certain regions or groups of people may not enjoy a similar quality of air and water as individuals in other regions.

[18] Carlino and Mills (1996) note that while state per capita incomes were converging throughout 1929–1990; they diverged from 1978 to 1988. In addition, most state income convergence occurred prior to 1946. Taken together, these facts strengthen our findings that the period surrounding 1970 characterized a structural break. For example, in spite of diverging per capita incomes in the post-1970 period, per capita emissions were converging in more states than before 1970.

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APPENDIX

An Example of Convergence and Divergence of Pollution

This appendix serves as an illustration of the point made in the text that, in theory, pollution levels may converge or diverge over time. We use a simple Solow growth model to make this point. Along a balanced growth path in such a model, the growth rate of per capita income $\hat{y}^*$and capital $\hat{k}^*$ equals the exogenous rate of technical change, g. When economies are off the balanced path, conditional convergence of incomes occurs

$$\text{(A1)} \qquad \hat{y} = g + \lambda(\ln y - \ln y^*),$$

where $\lambda = (\rho - 1)(\delta + n + g) < 0$. In (A1), y is per capita income, $\rho < 1$ is a production parameter, δ is the depreciation rate of capital, and n is the growth rate of population.

How does income convergence affect the pollution path over time? Assume that a social planner in state j maximizes utility defined as $U_{\mathrm{j}} = C_{\mathrm{j}} - zP_{\mathrm{j}}$, where C_{j} is aggregate consumption in state j, P_{j} is the level of pollution, and z is a preference parameter. Omitting subscripts, the budget restriction for each individual reads as follows: $y = c + b + i$, or $c + b = (1 - s)y$, where b is per capita abatement effort and s is the exogenous savings rate. Aggregate consumption is defined as per capita consumption c multiplied by the number of people in the state, or $C = L_0 e^{nt} c$ (L_0 is the initial population in the state). The state government chooses optimal per capita consumption c^* and abatement levels b^*, ignoring spillover costs affecting neighboring states.

Flow pollution is a byproduct of consumption but may be mitigated by abatement. In what follows, we adopt the specification by Andreoni and Levinson (2001)

$$\text{(A2)} \qquad P = C - C^{\alpha} B^{\beta},$$

where B represents aggregate abatement effort ($B = L_0 e^{nt} b$), so $C^{\alpha} B^{\beta}$ represents total abatement. Aggregate utility may be written as $U = C - z(C— C^{\alpha} B^{\beta})$, and upon setting z = 1, this simplifies to $U = C^{\alpha} B^{\beta}$. The assumption that $z = 1$ does not affect the qualitative nature of the results, but influences the optimal "level" of consumption and abatement. If the planner maximizes per capita utility at the state level, subject to the budget constraint, the optimal solution is defined as

$$c^* = \frac{\alpha}{\alpha+\beta} y(1-s), \text{ and} \tag{A3}$$

$$b^* = \frac{\beta}{\alpha+\beta} y(1-s). \tag{A4}$$

Upon substituting (A3) and (A4) into the pollution function and assuming that the economy is on a balanced growth path, the optimal quantity of pollution is a function of income, population, technical parameters (α, β), and other parameters (s, g)

$$P^*(y) = L_0 e^{nt}\left[\frac{\alpha}{\alpha+\beta} y_0^* e^{gt}(1-s) - \left(\frac{\alpha}{\alpha+\beta}\right)^{\alpha}\left(\frac{\beta}{\alpha+\beta}\right)^{\beta}(1-s)^{\alpha+\beta}(y_0^* e^{gt})^{\alpha+\beta}\right]. \tag{A5}$$

While this formulation is slightly different from the model developed by Andreoni and Levinson (specifically, both income and population are growing over time), the main findings are identical—assuming increasing returns to scale in abatement ($\alpha + \beta > 1$), the EKC obtains: $\partial P^*/\partial y^* \geq 0$ and $\partial^2 P^*/\partial y^{*2} \leq 0$. From (A5) we obtain three insights about the convergence of pollutants.

First, note that (A5) implies $P^* \to -\infty$ as $y(t) \to \infty$, which is not feasible. When adding $P \geq 0$ as an additional restriction, we can divide the dynamics of pollution into two phases. First, pollution levels are positive (an interior solution) and for sufficiently low initial income levels will describe an inverted-U shape over time. Later, pollution becomes zero (a corner solution) and remains zero thereafter. In the long run, absolute convergence of pollution is expected to occur at the zero pollution level.

Second, during the stage with an interior solution ($P > 0$) we can solve for the "balanced pollution path" as follows

$$\dot{P}^*(y) = nP + [\mu y - (\alpha+\beta)\theta(y)^{\alpha+\beta}]gL_0 e^{nt}, \tag{A6}$$

where $\mu = \frac{\alpha(1-s)}{\alpha+\beta} > 0$ and $\theta = (\frac{\alpha}{\alpha+\beta})^{\alpha}(\frac{\beta}{\alpha+\beta})^{\beta}(1-s)^{\alpha+\beta} > 0$. The rate of change of pollution along the interior balanced path is defined as

$$\hat{P}^* = \frac{\dot{P}^*}{P^*} = n + g\left[1 + \frac{(1-\alpha-\beta)\theta y^{\alpha+\beta}}{\mu y - \theta(y)^{\alpha+\beta}}\right]. \tag{A7}$$

With $\alpha + \beta > 1$, it follows from (A7) that the sign of the growth rate of pollution is ambiguous—for low-income levels the balanced pollution path is upward sloping, but as income increases the balanced pollution path levels off and eventually slopes downward (until it reaches the border where $P = 0$, after which $P = 0$ henceforth). Note that the rate of change of pollution is also a function of the "domestic" parameters, such as the propensity to save and production parameters.[19] These parameters affect the *level* of the balanced growth path of income, and the *rate of change* of the balanced pollution path.

Pollution levels are expected to converge conditionally to the balanced pollution path when economies are off the balanced income path

$$\hat{P} = (\hat{P}^{*})_{y=y*} + \frac{d\hat{P}}{dy}(y - y^{*}). \tag{A8}$$

To analyze out-of-equilibrium pollution levels we need to determine $\frac{d\hat{P}}{dy}$. Differentiating the expression for the pollution growth rate in (A7) gives us

$$\frac{d\hat{P}}{dy} = \frac{-g\theta(1-\alpha-\beta)^2\mu y^{\alpha+\beta}}{[\mu y - \theta(y)^{\alpha+\beta}]^2} < 0. \tag{A9}$$

The dynamics of out-of-equilibrium behavior of pollution may be summarized as

$$\hat{P} = (\hat{P}^{*})_{y=y*} + \frac{-g\theta\mu(1-\alpha-\beta)^2 y^{\alpha+\beta}}{[\mu y - \theta(y)^{\alpha+\beta}]^2}(y - y^{*}). \tag{A10}$$

We can now confirm that pollution levels converge towards the balanced path over time. If income is below (above) the balanced growth path, the growth rate of pollution exceeds (falls short of) the growth rate along the equilibrium pollution path.[20]

Third, and discussed in the main text, for interior solutions ($P > 0$), absolute pollution levels may converge or diverge depending on income levels—both during the approach to the balanced growth paths and along balanced growth paths.

[19]When the parameters α, β, s, n, and z take on different values for different states, income is expected to "conditionally" converge to a constant differential.

[20]The convergence process for pollution is slightly more complex than the convergence of income. Recall that the relationship between pollution and income is described by an inverted U for $\alpha+\beta>1$. Thus, the time path for pollution may be upward-or downward-sloping depending on the associated income level – hence the ambiguity in (A7). If the pollution path is downward-sloping, (A10) should be interpreted as follows: if actual income is lower than equilibrium income, (i) income will grow *faster* than along the balanced growth path, and (ii) pollution will be reduced *more slowly* than along the balanced growth path. If income exceeds income on the balanced growth path, (i) income will grow *more slowly* but (ii) pollution levels will be reduced *faster* than along the relevant equilibrium path. The pollution responses will be reversed if the economy is (still) located on the upward-sloping portion of the EKC.