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Measuring the effects of air quality regulations on "dirty" firm births: Evidence from the neo- and mature-regulatory periods*

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Abstract. In this article, we use annual (1980–90) county-level manufacturing plant location data for New York State to examine the effects of the 1977 Clean Air Act Amendments on the location decisions of new pollution-intensive manufacturing plants in the "neo-regulatory" (1980-84) and "mature-regulatory" (1985–90) phases of the Act's implementation. Our results suggest that the temporal effects of regulation vary. Whereas the location decisions of pollution intensive manufacturing firms were unaffected by the Act's regulatory restrictions in the "neo-regulatory" period, the restrictions appear to have had a significant negative impact on the location decisions of these types of firms in the Act's "mature-regulatory" phase. The diversion of new pollution intensive plants to counties with less stringent environmental regulations suggests that current US environmental regulations may be leading to a "browning process" whereby counties historically free of pollution become havens for polluters.

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1 Introduction

The impact of interjurisdictional competition on the structure of social regulations has sparked a controversial debate among academics, legislators, and the

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general public. Within this broad set of agendas, a growing interest in the effects of interjurisdictional competition on the basis of environmental regulations has recently developed. On the theoretical side, the bulk of the literature suggests that interjurisdictional competition will induce local regulators to “race to the bottom” and set suboptimal environmental standards to attract mobile capital (see e.g., Cumberland 1981; Markusen et al. 1995).

Although most theoretical models, as well as many national and local policy makers, assume the trade-off is clear, recent empirical evidence implies that stringency of environmental regulations is only weakly (or not at all) associated with the level of manufacturing activity (Bartik 1988; McConnell and Schwab 1990; Duffy-Deno 1992; Garafolo and Malhotra 1995; Levinson 1996b).[1] If these empirical findings do indeed represent reality, it is difficult to understand why so many state and local policy makers continue to provide tax relief and other financial incentives to help defray the costs of manufacturers’ compliance with pollution control regulations.[2]

While it would be easy to dismiss policy makers’ actions as irrational or self-serving behavior designed to garner attention and political support, the possibility that previous empirical research may have failed to address the problem at hand must be considered. We believe there are three interrelated empirical shortcomings that cast doubt on the previous empirical findings. First, due to data constraints, most location studies use aggregate manufacturing data. This is likely to lead to false inference – while the effects of environmental regulation may be minimal at the aggregate level, some industries, especially those that are heavy polluters or geographically mobile, may be significantly affected by stringent pollution control policies.[3] Second, the empirical model used by many researchers is cross-sectional and therefore does not control for the simultaneous nature of firm location and pollution problems. Since the location of industry, higher pollution levels, and subsequently more stringent pollution standards are positively correlated cross-sectionally, regression models that do not account for this phenomenon potentially yield inaccurate coefficient estimates.[4] In a related vein, a cross-sectional estimation procedure fails to address adequately the issue at hand. For example, analyzing the difference between new firm births in region i at time t and new firm births in region i at time $t + \phi$, as a function of the difference in region i’s environmental regulations over this same time period is

[1] Two exceptions are the recent county-level work of Henderson (1996), and the state-level study of List and Kunce (1999), which suggest pollution intensive industries are affected by more stringent pollution regulations.

[2] In its 1997 survey of state and local tax and financial incentives, Site Selection (1997) lists 15 instances in which states provided tax or financial incentives directly linked to investment in pollution control equipment by manufacturers. They also cite numerous other instances in which states provided tax and financial incentives that apply to equipment investments that are not restricted to, but could include, pollution control equipment.

[3] A few recent published exceptions are McConnell and Schwab (1990, automobiles), Henderson (1996, chemicals), and Levinson (1996b, various industries).

[4] See Henderson (1996) for a more complete discussion of this shortcoming.

much more informative than examining between region variation of these same variables at a snapshot in time.

In this article we use a unique panel data set to test whether county-level variations in environmental regulatory stringency, as measured by annual county-level attainment status of the primary federal standard for ozone, affect the location decisions of new pollution-intensive plants. The data set was constructed using individual manufacturing plant location data for New York State over the period 1980-90. Estimation results from a Poisson panel data model imply that the location decisions of these types of plants were affected differently in the early and later stages of the Act's implementation. Whereas the location decisions of dirty firms were unaffected by the Act's regulatory restrictions in the "neo-regulatory" period, the restrictions appear to have had a significant negative impact on the location decisions of these types of firms in the Act's "mature-regulatory" phase. The diversion of new pollution-intensive plants to counties with less stringent environmental regulations in the mature-regulatory period suggests that current US environmental regulations may be leading to a "browning process" whereby counties historically free of pollution become havens for polluters. Naturally, if the solution to air pollution is dilution, this particular altering of factor flows could represent an overall welfare improvement.

We organize the remainder of the article as follows. Section 2 provides a brief overview of air quality regulations in the United States. Section 3 describes the data, presents the empirical model, and discusses various econometric issues. Section 4 contains the empirical results, and a summary is provided in Sect. 5.

2 Air quality regulation

Prior to the early 1960s, responsibility for regulating air polluters in the United States rested almost exclusively with the states. Disappointed with the outcomes associated with decentralized control of the environment, federal authorities began to take a more active role in environmental regulation with passage of the National Environmental Policy Act and the first Clean Air Act Amendments (CAAA) in 1970. Federal organizations, such as the Environmental Protection Agency and the Council on Environmental Quality, were soon created to administer and enforce these statutes. As environmental authority was being shifted to Washington D.C., technological and political changes were making industry more geographically mobile, potentially intensifying and promoting inter-jurisdictional competition (see e.g., Tannenwald 1997). Coupling these temporal changes in firm mobility with the confusion of how the states should implement the new regulations of the CAAA of 1970, the federal government passed important provisioning rules via the 1977 Clean Air Act Amendments.

The 1977 Clean Air Act Amendments stipulated that starting in 1978 every county in the U.S. was to be designated annually as being in-attainment or out-of-attainment (non-attainment) of national air quality standards. A county's attainment status was to be determined with respect to each of five criteria air pollutants – carbon monoxide, sulfur dioxide, total suspended particulates, ozone, and nitrogen oxide (other pollutants, such as particulate matters, have subsequently been added to the list). If a county is not in attainment of the federal standard with respect to one of these pollutants, the state is required to submit periodic comprehensive plans that will lead to attainment status in the near future (commonly termed State Implementation Plans – SIP). If standards are not met in due time, states run the risk of losing federal monies that help to fund state-level public goods and services.

Environmental regulations in non-attainment counties can potentially be very costly for new plants. Firms entering a county labeled out-of-attainment are subject to a standard of "Lowest Achievable Emission Rate [LAER]" on equipment, without consideration of cost. These abatement expenditures potentially run into the millions of dollars and represent a significant set-up cost for firms in pollution intensive sectors. New plants locating in attainment areas, on the other hand, face a more lax regulatory standard. These plants are subject to the standard of "Prevention of Significant Deterioration [PSD]". This entails permitting and the installation of the "Best Available Control Technology [BACT]" for new plants that have the potential to emit over 100 tons of a criteria pollutant in a year. The BACT is negotiated on a case by case basis and the economic burden on the plant is taken into consideration in arriving at a final solution. Given that the installation of BACT in attainment areas is likely to be much less costly than the installation of LAER in non-attainment areas, new polluting plants could face significantly lower pollution control capital construction costs in attainment areas versus non-attainment counties.

3 The data and the empirical model

3.1 The data

The plant location data for this study is comprised of a set of 282 individual manufacturing plants that opened in New York State during the period 1980–1990. We construct the data file from the comprehensive Industrial Migration File (IMF) that was maintained until 1990 by the New York State Department of Economic Development (DED). The intent of the IMF was to monitor all gains and losses in manufacturing activity in NYS by county on an annual basis. The data units in the IMF are case observations of individual plant openings, closings, expansions, and contractions. The information in the IMF file was assembled from a variety of sources, including regional offices of the NYS Department of Commerce, local chambers of commerce, the NYS Department of Labor, newsarticles, and private reporting sources. The DED regional offices verified all reported projects before including them in the file. This method of data collection has both advantages and disadvantages.

A major advantage is that case-specific information is not suppressed (Michalke 1986). Thus, there are opportunities for improving precision and data analysis not heretofore available even from the Longitudinal Research Database (McGuckin 1990). On the negative side, since no legislative statute exists in

New York that requires plants to furnish information to the state, DED makes no claim that the IMF is comprehensive. In addition, the IMF excludes some plant activities involving either small investment activity (less than $ 100,000) and/or modest changes in employment (less than 25 employees). Nevertheless, comparisons with Census of Manufacturers data suggest that IMF coverage is extremely broad for all but the smallest size classes.

3.2 Pollution intensive industries

The first step in preparing our plant location data for analysis was to designate major SIC industry codes as pollution intensive. Many facets of the complex pollution process need to be considered when grouping industries into a polluting set. We follow List and Co (1999) and use three pieces of information to designate 2-digit SIC industries as pollution intensive: 1) firm-level pollution abatement operating expenditures to abate the media air, water, and solid/contained waste from the Current Industrial Report’s *Pollution Abatement Costs and Expenditures* database; 2) actual emissions, by industry, of the five criteria air pollutants from 1980 and 1990 from the *National Air Pollutant Emission Trends* 1900–1994, published by the Environmental Protection Agency (EPA); and 3) the criteria established in the pollution literature (e.g., Levinson 1996b and Jaffe et al. 1994). Based on this troika, we classified seven 2-digit SIC industry codes (26, 28, 29, 32, 33, 34, and 37) as pollution intensive, and used data from these plant types in our analysis.

The number of firm openings by year and 2-digit SIC industry codes are shown in Table 1. In general, the total number of new firm openings shows a slight downward trend over the decade. The last column in Table 1 shows the percentage of firms in each industry that located in counties that were in attainment of the federal primary standard for ozone. The percentages suggest a preference for location in attainment counties. The percentages of firms that chose a site in an attainment county group ranges from a low of 47% to a high of 85% with an overall percentage of 60%. These percentages are consistent with a *priori* expectations and suggest a graying process may be occurring as polluting firms prefer to locate in low stringency counties. The results in Table 1 should be considered unconditional, however, since there has been no attempt to control for other factors that affect the plant-siting process. These other factors, which range from county-level amenities to employee wage rates, can be adequately accounted for in a well-specified econometric model.

3.3 The empirical model

To derive our reduced-form empirical model, we follow Becker and Henderson (1997) and revise the stock construct in Henderson et al. (1995) to a flow concept. This general econometric approach enables us to capture the effect of more stringent pollution standards in a partial equilibrium framework. We assume that

Table 1. Count of firm openings in New York State by year and industry

Pollution intensive industries													
SIC Code	1980	1981	1982	1983	1984	1985	1986	1987	1988	1989	1990	Industry total	%in attainment counties
26	5	3	5	4	11	5	9	6	2	2	2	54	57%
28	10	3	2	4	3	7	4	5	2	1	5	46	57%
29	1	0	0	0	0	1	1	1	0	0	0	4	75%
32	1	4	2	1	1	2	1	0	3	3	2	20	85%
33	7	2	5	4	1	2	4	7	5	2	2	41	68%
34	13	9	10	10	10	5	4	9	8	5	4	87	47%
37	3	1	3	3	4	4	3	3	3	1	2	30	80%
Total	40	22	27	26	30	26	26	31	23	14	17	282	60%

in any given time period there is a supply of entrepreneurs in each county that is given by $Y_{it}(\prod(X_{it}, e_{it}))$; where Y_{it} represents the flow of new plants, X_{it} are spatial attributes that affect the local profit function, Π, and e_{it} is a random error component. If one envisioned this curve in a Cartesian plane of births and per plant profit space, it would be upward sloping, indicating that an increase in per firm profits is necessary to induce the marginal entrepreneur to start-up a firm. The X_{it} are therefore shift parameters that influence the location of the curve.

Completing the picture, one can consider profit opportunities for new plants as being represented by a "demand curve" that depicts changes in per plant profits as a result of an increase in births, $\Pi^d(Y_{it})$. Since births can positively affect per plant profits through localization and/or urbanization economies, or detract from per plant profits through a negative impact on local output price, the demand curve can slope upward or downward. Given that equilibrium occurs at the intersection of the demand and supply functions, we equate and solve for Y_{it} to get the reduced-form equation for a continuous event:

$$Y_{it} = f(X_{it}) + e_{it} \tag{1}$$

where Y_{it} is the count of new plants in county i at time t; X_{it} includes county attributes presumed to affect the spatial location function, and e_{it} is the well-behaved contemporaneous error term.

Because Y_{it} are strictly integer values that are generally close to zero, we model the regressand as a Poisson distributed random variable. The Poisson distribution models the discrete nature of the dependent variable and naturally models zero outcomes. Because unobservable county-specific factors that may affect the location decision, α_i, are potentially important, we estimate the Poisson panel data model of Hausman et al. (1984; hereafter HHG). The basic Poisson probability function is expressed as:

$$\text{Prob}(Y_{it}) = \frac{e^{-\lambda_{it}}\lambda_{it}^{y_{it}}}{Y_{it}!}, \qquad y = 0, 11, \ldots; \tag{2}$$

where Y_{it} denotes new plant births in county i at time t, and λ_{it} is the Poisson parameter. As in the HHG model, the Poisson parameter is given by:

$$\lambda_{it} = e^{x_{it}\beta + \alpha_i}, \tag{3}$$

where $\ln \lambda_{it} = \beta' X_{it} + \alpha_i$. X_{it} and α_i are defined above, and β is the vector of unknown parameters to be estimated. As such, the Poisson parameter, λ_{it}, is the expected value of new births. The property of the Poisson distribution indicates that the expected value of Y_{it} is equivalent to the variance of Y_{it}. It is not unusual, however, that the variance is larger than the expected value, implying "overdispersion" in the empirical data.

Although the natural approach to account for heterogeneity in the standard linear panel data model is to transform the variables about their means, this does not remove the heterogeneity in the Poisson framework. Instead, the fixed effects are accounted for by modeling the within county sequence of births, conditioned on total births for that county over the sampling period. Hence, the likelihood function becomes:

$$\text{Prob}\left(Y_{i1}, Y_{i2}, \ldots Y_{iT} \,\Big|\, \sum_t Y_{it}\right) = \frac{\left(\sum_t Y_{it}\right)!}{\Pi_t\,(Y_{it}!)}\,\Pi_t \left[e^{x_{it}\beta} \Big/ \sum_t e^{x_{it}\beta}\right]^{Y_{it}} \tag{4}$$

Since the fixed effects, α_i, have been conditioned out of the likelihood function they are unrecoverable, and indeed not estimated.[5]

A few noteworthy aspects of our estimation procedure merit further attention. First, estimation of the HHG model places the same demands on the data as the typical linear fixed effects panel data model: between-county variation is captured by the county fixed effects, leaving only within-county variation to be explained. An entrepreneur scanning the entire horizon and choosing the profit maximizing location represents an alternative modeling procedure. A model consistent with this assumption is the conditional logit framework of Levinson (1996a) and List and Co (1999). The modeling choice represents an interesting trade-off that is oftentimes neglected in the literature. If one attempts to model the attractiveness of a county relative to other counties, a cross-sectional conditional logit model may be appropriate, as between county variation is captured by the response coefficients. Alternatively, if one conjectures that unobserved characteristics of counties (such as labor-force worker ethic) are important in the siting decision, and controlling for these factors is invaluable, a within-county approach similar to ours is appropriate. We believe that analyzing the difference between new firm births in county i at time t and new firm births in county i at time $t + \phi$, as a function of the difference in county i's environmental regulations over this same time period is much more informative than examining between county variation of these same variables at a snapshot in time. Our within-county approach avoids

[5] Our partial equilibrium model is similar in nature to the counting approach used in List (1999). Also, note that our likelihood function is (up to a constant) equal to the likelihood function of the multinomial logit model.

many of the pitfalls, such as omitted variable bias, that are inherent in cross-sectional models.

A second important characteristic of our Poisson framework is that λ_{it} is assumed to be both the mean and variance of Y_{it}. Failure of this restriction has consequences similar to those of heteroscedasticity in the ordinary least squares framework: parameter estimates are consistent, but variances are inconsistently estimated, leading to invalid hypothesis tests. We use Wooldridge (1991) robust standard errors – robust to violation of the moment restriction – to elude this restriction. The Wooldridge standard errors are robust to both variance misspecification and arbitrary serial correlation, and therefore they tend to be larger than standard errors from the Poisson model when there is overdispersion in the data. Third, the estimated coefficients from the maximization of equation (4) are consistent only if the conditional mean is correctly specified and the true distribution is linear exponential (need not be Poisson). In each set of estimates, we include a robust Hausman test for model misspecification, which compares the HHG estimates with consistent estimates from non-linear least squares estimation with the same conditional mean function.

With respect to the regressors our primary focus in vector X_{it} is the county-level attainment variable. As previously mentioned, the 1977 CAAA set standards on five criteria air pollutants – sulfur oxides (SO_x), carbon monoxides (CO), ozone (O_3), nitrogen oxides (NO_x), and total suspended particulates (TSP). Because ozone has attracted the most regulatory attention due to the limited progress that has been made to reduce concentration levels, we follow Henderson (1996, 1997) and focus on county attainment status of ozone. Although attainment status can range from attainment of the primary standard to non-attainment, with partial standards in between, ozone designation has essentially been polar in nature – that is a county is either in, or out, of attainment. We therefore construct a dichotomous variable that is equal to 1 if the county is out-of-attainment of the primary federal standard for ozone and 0 otherwise.

A further issue regarding the effect of attainment status is that there may have been a qualitative difference in regulation enforcement in the period immediately following the implementation of the 1977 CAAA compared to later periods. A well- known fact is that in the early regulatory years many forms of production activity that are now considered hazardous were not deemed detrimental by local and federal regulators. In addition, at the onset of the 1977 CAAA, many small emitters of pollutants were largely ignored to allow a more focused effort on major, class A emitters. Consequently, the early regulatory era may have witnessed much different location patterns compared to later years.[6] In maximizing the likelihood function in equation (4), we explicitly allow for temporal differences in attainment status effects.

Besides the annual air pollution regulatory variable, we include in vector X_{it} traditional exogenous variables that are common arguments in the firm's location function. These control variables represent proxies for labor market

[6] See Becker and Henderson (1997) for a more complete discussion of this point.

characteristics, market size, and the tax climate in a particular county. We use annual data on wage rates and population to control for the first two factors. Within our sample period, property tax data appear to be the best candidate to control for the county's tax climate. Unfortunately, these data are only available for 1982 and 1987. We therefore use 1982 property tax data for years 1980–1987 and 1987 data for 1987–1990. We initially used 1977 data for 1980–1982, but results were largely unchanged. Since we are using this regressor for control purposes only, we are not too concerned with its lack of variation. Descriptive statistics and definitions of all variables are presented in Table 2.

Table 2. Description of variables[a,b]

Variable	Mean (Std. Dev.)	Definition and source
Pollution intensive New plants	0.41 (1.0)	Actual count of new plants entering the county from 1980–1990. Industrial Migration File. New York State Dept. of Economic Development (DED). Firms labeled as having production activities that are pollution intensive. SIC codes 26, 28, 29, 32, 33, 34, 37.
Attainment Status	0.28 (0.45)	Intensity of county-level pollution regulations. Dichotomous variable = 1 if county is out-of-attainment of federal standards for ozone, 0 otherwise. Federal register Title 40 CFR Part 81.305.
ln wage	9.71 (0.23)	Natural logarithm of total annual manufacturing payroll divided by the number the number of employees by county, adjusted for inflation. *County Business Patterns.*
ln population	11.66 (1.25)	Natural logarithm of county population. *Current Population Reports.* U.S. Bureau of Census.
Property tax	562.37 (248.03)	Property tax collected per capita, adjusted for inflation. *Census of Governments.*

[a] Data are for the 62 New York counties.
[b] Data for property taxes are only available for 1982 and 1987. The remaining variables are annual.

4 Empirical results

Table 3 contains estimation results for two specifications of our data. Before discussing the coefficient estimates, four preliminary remarks are useful. First, in each model type, we initially allowed attainment status to have an heterogeneous effect on annual location patterns. Since early years of regulation could be thought of as a much different regime than later years, we proceeded to split the sample into a "neo-regulatory" period, 1980–1984, and a "mature regulatory" period 1985–1990. Likelihood ratio tests suggest that this grouping scheme is not unduly

Table 3. Conditional poisson panel data estimates of the determinants of county-level firm location

Independent variable	(1)	(2)	Independent variable	(1)	(2)
Attainment Status (1980–84)	−0.17 (0.26)	−0.23 (0.27)	1981	−0.22 (0.29)	−0.15 (0.33)
Attainment Status (1985–90)	−0.88 (0.45)	−0.95 (0.44)	1982	0.32 (0.34)	0.44 (0.42)
Robust-ttest of slope differences	−1.56	−1.63	1983	0.55 (0.38)	0.70 (0.51)
Wage	−4.79 (2.44)	−5.58 (3.00)	1984	0.95 (0.60)	1.12 (0.73)
Population	–	4.66 (3.36)	1985	1.10 (0.74)	1.29 (0.88)
Property tax	–	−0.001 (0.002)	1986	1.23 (0.86)	1.45 (1.04)
Robust Hausman statistic	14.31	17.97	1987	1.63 (0.97)	2.03 (0.97)
R^2	0.57	0.57	1988	1.58 (1.10)	1.97 (1.13)
N	682	682	1989	1.32 (1.18)	1.74 (1.21)
			1990	1.75 (1.33)	2.16 (1.35)

Notes:

1. Dependent variable is the annual count of new plants from 1980–1990. Attainment status and population are measured annually; remaining regressors are measured in 1982 and 1987.
2. Robust standard errors are in parentheses beneath coefficient estimates.
3. Coefficient estimates are marginal effects computed at the sample means.
4. Robust Hausman statistic is a test of model significance.
5. Fixed county-effects are included in each regression model, and are jointly significant at the 1% level.

restrictive – in both specifications we could not statistically distinguish between this pooling procedure and the totally flexible model that allows each year to have a different effect.

Second, we estimate a relatively parsimonious model, akin to Henderson's (1996, 1997) formulations, and a less stark version which includes other county attributes that a new firm may consider important in the siting decision. Parameter estimates suggest that the stark model, which includes only attainment status, wage rates, and time and county dummies as regressors, performs similarly to models that also include property tax and population as regressors. This lends credence to Henderson's approach as many of the factors that affect plant births, such as land availability, worker ethic, and macroeconomic influences, are captured by the county and time-specific dummy variables. Nonetheless, both models are significant at the $p < 0.01$ level.

Third, a goodness of fit measure and specification tests both suggest that our Poisson model is an appropriate estimation procedure. For the former, we calculate a measure of fit, R^2 = 1 – (residual sum of squares/total sum of squares), or more appropriately:

$$R^2 = 1 - \left(\sum_T \sum_N \left(Y_{jt} - E\left(Y_{jt}\right)\right)^2 \right) \Big/ \left(\sum_T \sum_N \left(Y_{jt} - \overline{Y}_N \right)^2 \right). \qquad (5)$$

Estimated R^2 values in the bottom panel of Table 3 are nearly 0.60, implying that our relatively parsimonious models perform quite well. In terms of model misspecification, our diagnostic procedures imply that our fixed effects Poisson model is appropriate. In each specification, the regression based robust Hausman statistic in the lower portion of Table 3 is less than the critical values for the χ^2 statistic at the 95% percent level – 22.4 for the stark model and 25.0 for the less stark model – suggesting the null of proper specification cannot be rejected for either of the models.

Fourth, endogeneity of our attainment status variable may be a potentially serious issue. Since firms entering a county may increase overall pollution levels, inducing a county to be labeled out-of-attainment of federal air quality standards, endogeneity could cause a downward bias on the coefficient of attainment status. We follow Papke (1991) and empirically check for endogeneity by comparing results from models including contemporaneous attainment status with those using lagged values of attainment status. In each case the coefficients do not change markedly, suggesting endogeneity is not a major problem inherent in our data. We therefore present results from the contemporaneous specifications. Finally, since the coefficient estimates of our environmental measure rely on attainment status changes within a county, it is important to note that 29 of the 62 counties had attainment status changes during our sample period.

Coefficient estimates in Table 3 provide many interesting insights. For example, the estimated time-effects in the right portion of Table 3 suggest that, controlling for county effects and other county attributes, there was an upward trend in new firm births in New York State from 1980–1990. More importantly, response coefficients in columns 1 and 2 imply that "dirty" firms were paying significant attention to county attainment status, particularly in the "mature regulatory" period. Although in the neo-regulatory period attainment status has an insignificant effect on location patterns, estimates in the stark (non-stark) model suggest that when a county changed from in-attainment to out-of-attainment, its expected flow of births decreased by 0.88 (0.95) between 1985–1990. Both coefficient estimates are significantly different from zero at the $p < 0.05$ level, and suggest that attainment status significantly effects the annual expected flow of new firm births for the average county from 1985–1990.

Robust t-tests of slope differences suggest that these effects are larger in magnitude than parameter estimates in the neo-regulatory period at the $p < 0.06$ significance level (one tailed $t = -1.56$; $t = -1.63$). This finding suggests that regulation was qualitatively different throughout the 1980s. Whereas in the later years regulators began to recognize and enforce statutes against even small and medium emitters, the early regulatory years were characterized by small emitters of pollutants being largely ignored to allow a more focused effort on major, class A emitters. Consequently, the early regulatory era had less stringently imposed standards. This finding may explain why the majority of empirical studies that use data from the 1970s and early 1980s – eras of relatively small regulatory compliance expenditures compared to other firm outlays – find insignificant effects of stricter environmental regulations.

Although larger in magnitude, these effects are in the neighborhood of Henderson's (1997) estimates for the chemicals industry. The results suggest that after the initial administration issues were settled in the neo-regulator era, federal air pollution regulations induced a browning process as polluting firms were more likely to choose counties with relatively cleaner environments and lower regulation costs. In light of Wasylenko (1997), these larger magnitudes are intuitively plausible since smaller areas tend to have less variation in other important location factors, such as labor markets, climate, work ethic, amenities, and energy costs, which in turn accentuates any differences in attainment status.[7]

Other parameter estimates largely have expected signs. The estimated coefficient of the wage variable in each model suggests that an increase in wage schedules leads to a significant decrease in the expected flow of new firm births in the average county. These estimates are larger than comparable findings in state-level studies, and may be a repercussion of using a smaller area of study, which tends to accentuate variation in wage rates. Other coefficients in the model indicate that higher property taxes deter new firm births, but not significantly. On the demand side, increases in population attract new firms, but not at statistically significant levels. This unexpected finding is probably an indication that the county fixed effects are capturing many of the demand side considerations.

[7] To put our results into perspective, we advise the interested reader to see Tannenwald (1997), Gerking and List (1999), and Levinson (1996a), who provide a catalogue of previous empirical studies that focus on the consequences of more stringent environmental regulations. Also, note that some of these studies use another technique surveys to measure the effects of environmental regulation on economic activity. Given the lack of incentive compatibility that plagues hypothetical surveys (see e.g., List and Shogren 1998), we avoid this procedure.

5 Conclusions

Do environmental regulations alter factor flows? This question remains an important public policy issue as policymakers and the popular press both suggest that the question has a simple affirmative answer. This reasoning adheres to basic intuition as environmental regulations represent a constraint on the economic system, which forces adaptations and a redistribution of capital and people. Even so, a majority of the extant empirical evidence in academic circles suggests that the effects are minimal or non-existent. We use a new panel data set from 1980–1990 to provide further intuition into this puzzling gap between the empirical evidence and most people's intuition. Our results suggest that manufacturers in pollution intensive sectors are deterred severely by more stringent county-level environmental regulations, particularly after the EPA and state agencies had an opportunity to experiment and develop a systematic procedure to deal with polluters.

These results suggest that a temporal browning process may be occurring across space. Whether the writers of federal air quality regulation intended this or not, our evidence suggests that environmental regulations alter factor movements. Nonetheless, certain factor flows may represent a Pareto improvement as areas that have abundant natural resources are now attracting firms and those counties with depleted ecosystems are essentially given a "break" from further destructive manufacturing activity. We only state this as a matter of conjecture and note that more research on the welfare implications of this graying process needs to be carried out before further evaluating the efficacy of current environmental policies.

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