NBER WORKING PAPER SERIES

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Working Paper 7369
http://www.nber.org/papers/w7369

NATIONAL BUREAU OF ECONOMIC RESEARCH
1050 Massachusetts Avenue
Cambridge, MA 02138
September 1999

This paper was first presented as part of the *International Seminar on International Trade* (ISIT) conference. We have benefitted from funding provided by the National Science Foundation and the Wisconsin Alumni Research Foundation, and from invaluable research assistance by Joe Hendrickson. Wayne Gray, Kevin Hassett, Jim Hines, John List, Gib Metcalf, Domenico Siniscalco, Matthew Slaughter, and numerous participants in seminars at the ISIT conference andthe NBER have provided especially constructive comments. Any opinions expressed are those of the authors, and not of the National Bureau of Economic Research.

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NBER Working Paper No. 7369
September 1999
JEL No. F2, H73, Q2

ABSTRACT

This paper estimates the extent to which changing environmental standards have altered patterns of international investment. Our analysis goes beyond the existing literature in three ways. First, we avoid comparing regulations in different *countries* by using data on inward foreign direct investment (FDI) to the U.S. and on differences in the regulatory stringency of U.S. *states*. This approach has the advantage that data on environmental stringency in U.S. states are more comparable than that for different countries, and that U.S. states are more similar than countries in other difficult-to-measure dimensions. Second, our measure of environmental stringency accounts for differences in states' industrial compositions, an acknowledged problem for earlier studies. Third, we employ a panel of annual measures of relative regulatory stringency from 1977 to 1994, allowing us to control for unobserved state characteristics that may be correlated with both FDI and compliance costs. We find some evidence of small deterrent effects of environmental regulations in particularly pollution-intensive industries, but no evidence of large or widespread effects. While the broad conclusions are consistent with the existing literature, this paper does address three important concerns with that literature.

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I. Introduction.

In recent years, a variety of interest groups have called for addenda to international trade agreements to harmonize domestic labor standards, antitrust policies, and environmental regulations. Industry representatives in the U.S. worry that stricter standards will put U.S. manufacturers at a competitive disadvantage. Environmentalists fear that linked trade agreements will prevent countries from setting their desired levels of environmental regulation. Free trade advocates worry that countries may be able to circumvent international agreement on tariffs by choosing strategic levels of domestic regulation (Ederington, 1998; Copeland, 1990). And some economists have worried that governments may seek to attract foreign direct investment (FDI) by competitively undercutting each other's environmental standards (Bhagwati and Srinivasan, 1995). All these fears are based on the presumption that domestic regulations affect the location of FDI in quantitatively important ways. This paper tests that presumption by asking whether FDI to U.S. states has responded significantly to relative changes in states' environmental compliance costs.

As recently as the Uruguay Round of the General Agreement on Tariffs and Trade negotiations, FDI was considered a "new issue" (Baldwin 1995). Today, however, the relationship between trade and FDI has become recognized, in part because multinational corporations associated with FDI account for a large and growing portion of world trade.[1] At a simple level, two processes work in opposite directions. As world trade barriers fall, firms may

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[1]In 1994, 36.3 percent of U.S. imports and 42.7 percent of U.S. exports involved intra-firm trade between multinational parents and affiliates (BEA, 1997).

be more easily able to outsource parts of their manufacturing operations, resulting in increasing FDI. On the other hand, trade agreements diminish one important motive for FDI--to circumvent tariffs.

From 1982 to 1996, the fraction of U.S.-owned manufacturing located outside U.S. borders rose from 10.9 percent to 13.0 percent, and overall U.S. FDI abroad rose from $224 to $498 billion, an average annual growth rate of 5.7 percent.[2] Some analysts have concluded that U.S. producers are taking advantage of falling tariffs by investing abroad to avoid relatively high taxes, factor costs, and regulatory standards in the U.S.[3] However, the fact that inward FDI to the U.S. has also increased over the last two decades suggests that there must be other forces working in favor of investing in the U.S.[4]

The specific issue we address in this paper is the implicit assumption that underlies calls for environmental harmonization: that FDI does in fact respond significantly to international differences in regulatory stringency. Despite numerous attempts in the economics literature, there is little robust or quantitatively significant evidence that environmental regulations affect

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[2]These numbers are for the GDP of majority-owned U.S. subsidiaries, from Bureau of Economic Analysis (BEA) data; this is the preferred measure of the economic importance of FDI, see Lipsey et al. (1998). Transactions are classified as FDI if a foreign entity owns 10 percent or more of the securities of an incorporated business, or an equivalent interest in an unincorporated business. Mataloni (1999) shows that for 80 to 85 percent of such foreign-owned manufacturers, the foreign entity owns more than a 50 percent interest.

[3]The overview of the evidence in Caves (1996) indicates that different production costs affect the location of FDI, and recent theoretical work by Horstmann and Markusen (1992), Brainard (1993), and Keller (1998) builds on this fact.

[4]Foreign-owned firms accounted for 2.6 percent of manufacturing GDP in the U.S. in 1977, and for 9.2 percent in 1996 (BEA 1999).

the location of FDI. The next section describes our empirical methodology and how it compares to previous analyses of environmental regulations and FDI, and section III presents our results.

II. Measuring the Effects of Regulations on FDI.

Most papers in this literature note the inherent difficulties in quantifying the stringency of national environmental standards. Even if one could accurately measure stringency, countries differ on so many other grounds that it is hard to attribute any differences in FDI inflows to environmental regulations. Until recently, most analysts have thus resorted to comparing investment in developing countries to that in industrialized countries, assuming that industrialized countries have more stringent standards (Low, 1992; Leonard, 1988). While this assumption seems realistic, the fact that industrialized countries are nevertheless the largest exporters of polluting goods suggests that differences in economic activity are not caused by environmental policy alone. World trading patterns are in part determined by factors and technologies that are not readily observable, and therefore difficult to control for statistically, and the same is likely true for FDI patterns.

We overcome the difficulties of comparing different countries by looking at the flow of investment from foreign countries into U.S. states as a function of state regulatory stringency. This gives us two advantages: there are much better data on state environmental compliance costs than on international costs, and different states are more comparable than different countries on non-environmental grounds. The states hold a large and increasing fraction of the responsibility for setting environmental standards in the U.S., and even those standards that are set federally impose different costs depending on the characteristics of the affected states.

We examine two types of FDI data. The first is data on the value of gross property, plant and equipment owned by foreign-owned manufacturers, and manufacturing employees working for foreign-owned firms, from the series *Foreign Direct Investment in the United States* of the Bureau of Economic Analysis (BEA). Though comprehensive, these data may have two disadvantages for our purposes. First, they include both new and existing facilities. Since most state environmental regulations impose stricter standards on new facilities, states with more new investment will have *higher* average compliance costs, which might induce a bias in our study against finding a deterrent effect of environmental regulations.

Second, the BEA data include mergers and acquisitions. If the regulatory differences among states are capitalized into purchase prices (foreign investors receive a discount when buying manufacturers in stringent states), then we would expect there to be no deterrent effect of strict regulations on mergers and acquisitions. Therefore, to avoid bias caused by differential treatment of new investment or compliance cost capitalization, as a second approach we examine planned new foreign-owned factory openings using data from a different series, also titled *Foreign Direct Investment in the United States*, collected by the International Trade Administration (ITA). These ITA data have the drawback that relatively few new foreign-owned manufacturing plants are observed in any given state in any year. From 1977 to 1994, the data contain only 958 new plants. Nevertheless, the ITA and BEA data together provide a comprehensive picture of FDI to U.S. states. By comparing foreign direct investment to different states rather than to different countries, we believe that we increase enormously our chances of accurately measuring regulatory stringency and of sufficiently controlling for other characteristics that attract or deter investment.

The second problem with the existing literature on the effects of environmental regulations is that most papers rely on cross-sectional data.[5] This makes it impossible to account for unobservable state characteristics that may be correlated with both regulatory stringency and investment. For example, suppose that some state is endowed with a natural resource desirable to a polluting industry. As a consequence, that state will be likely to attract investment in that industry, and may simultaneously be induced to regulate stringently the pollution emanating from the industry. Both investment and regulatory stringency will be *positively* correlated with the presence of the desirable resource, inducing a spurious positive correlation between FDI and environmental stringency. As another example, suppose that some states have a tendency to favor polluting industries, perhaps because those industries are particularly important to the states' economies, or because those industries have long histories in the states. In this case, investment and regulatory stringency will be *negatively* correlated. If the estimation does not account for the unobserved resource, or the unobserved protection of polluting industries, then it will impart an omitted variable bias on the predicted effect of regulatory stringency on investment.

By contrast, several recent studies of domestic investment use panel data and find reasonably sized and statistically significant negative effects of environmental stringency on economic activity. Henderson (1996), Greenstone (1998), and Kahn (1997) use data on whether or not each county in the U.S. is in compliance with national ambient air quality standards. These standards are set uniformly at the federal level, and are thus unrelated to particular county characteristics, whether observed or otherwise. States are required to enforce more stringent

[5]See, for example, Friedman, *et al.* (1992), Kolstad and Xing (1997), or Co and List (forthcoming1998).

pollution standards in counties declared out of compliance, and all three studies find that such counties subsequently experience fewer new plant births or less manufacturing employment growth. However, it is difficult to interpret the general magnitude of the effect of this zero-one measure of regulatory stringency without knowing how much more costly are the environmental regulations in non-compliant counties.

We address this second problem, omitted variable bias, by examining investment and regulatory stringency over an 18-year period, from 1977 to 1994, which allows us to control for unobserved time-invariant state characteristics in the estimations. Rather than use a zero-one measure of regulatory stringency, such as counties' compliance status, we use a continuous, time-varying measure of the pollution abatement costs in each state, based on data from the Pollution Abatement Costs and Expenditures (PACE) survey, conducted by the U.S. Census Bureau as part of the Annual Survey of Manufactures.

The third shortcoming of much of the existing literature on investment responses to environmental regulations is that quantitative measures of regulatory stringency typically fail to account for regions' industrial compositions. Friedman, *et al*. (1992), Crandall (1993), and Co and List (forthcoming), for example, measure environmental stringency using total statewide pollution abatement costs from the PACE survey, and conclude that investment is largely unaffected by environmental regulations. As they note, however, the problem with their measure of costs is that states that attract polluting industries will have higher abatement expenditures than states that have cleaner industrial compositions even if the regulatory stringency faced by individual firms is the same for all states.[6] If lax regulations do attract polluting industries, pollution abatement spending may in fact be negatively correlated with the stringency of state regulations.

We address this third problem by measuring state pollution abatement costs from the PACE data, adjusted using each state's industrial composition. Ideally, one would study this issue industry-by-industry, using separate measures of pollution abatement costs for each industry to assess regulatory stringency. While abatement costs by state and industry are published annually by the Census Bureau, so many of the observations are censored to prevent disclosure of confidential information that the data are not comparable year-to-year or state-to-state. Therefore, we propose an alternative index.[7]

The index compares the *actual* pollution abatement costs in each state, unadjusted for industrial composition, to the *predicted* abatement costs in each state, where the predictions are based solely on nationwide abatement expenditures by industry and each state's industrial composition.[8] Let the actual costs per dollar of output be denoted

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[6]Co and List (forthcoming) also examine inward FDI's cross-sectional correlation with state environmental agencies' budgets, and with ambient pollution readings in each state, with similar outcomes: coefficients are small, often statistically insignificant, and are not larger in magnitude for more pollution-intensive industries.

[7]More details about this index, and a comparison of it with other indices of state environmental standard stringency can be found in Levinson (1999). Gray (1998) and Levinson (1996) construct similar indices using the confidential plant-level Census data. The advantage of the index used here is that it is available publicly and yields information similar to that from the unpublished Census data.

[8]For two reasons, we use pollution abatement *operating expenses* (as opposed to *capital expenses*) in the index. First, operating expenses for pollution abatement equipment are easier for PACE survey respondents to identify separately. Abatement capital expenses may be difficult to disentangle from investments in production process changes that have little to do with pollution

$$S_{st} = \frac{P_{st}}{Y_{st}} \qquad \textbf{(1)}$$

where P_{st} is pollution abatement costs in state s in year t, and Y_{st} is the manufacturing sector's contribution to the gross state product (GSP) of state s in year t. S_{st} is the type of unadjusted measure of regulatory stringency commonly used, and it overstates the stringency of states with more pollution-intensive industries and understates the stringency of states with relatively clean industries.

To adjust for states' industrial compositions, compare **(1)** to the *predicted* abatement costs per dollar of GSP in state s:

$$\hat{S}_{st} = \frac{1}{Y_{st}} \sum_{i=20}^{39} \frac{Y_{sit}\ P_{it}}{Y_{it}} \quad , \qquad \textbf{(2)}$$

where industries are indexed from 20 through 39 according to their 2-digit manufacturing SIC codes,[9] Y_{sit} is industry i's contribution to the GSP of state s at time t, Y_{it} is the nationwide contribution of industry i to national gross domestic product, and P_{it} is the nationwide pollution abatement operating costs of industry i. In other words, $\hat{S}_{st}$ is the weighted average of national pollution abatement costs in each 2-digit industry, where the weights are the shares of each industry in state s at time t.

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abatement. Second, abatement capital expenditures are highest when new investment takes place. So states that have thriving economies and are generating manufacturing investment tend to have high levels of abatement capital expenses, regardless of the stringency of those states' environmental laws. Operating costs are more consistent year-to-year.

[9]SIC code 23 (apparel) is omitted because it is relatively pollution-free, and as a result no data for that industry are collected by the PACE survey.

The industry-adjusted index of relative state stringency, S_{st}*, is simply the ratio of actual expenditures in **(1)** to the predicted expenditures in **(2)**[10]

$$S_{st}^{*} = \frac{S_{st}}{\hat{S}_{st}} \quad . \tag{3}$$

When S_{st}* is greater than 1, industries in state *s* at time *t* spent more on pollution abatement than those same industries in other states. By implication, states with large values of S_{st}* have relatively more stringent regulations.[11]

In section III.2, we use the BEA's continuous measures of FDI to estimate models of three different types: a pooled ordinary least squares specification as a benchmark, a fixed-effects least-squares (within groups) estimator, and a dynamic panel data model that includes the lagged dependent variable as a regressor. In section III.3, we employ the ITA's data on new factory openings to estimate Poisson and other count data models. Before that, however, we begin with simple descriptive statistics.

III. The Evidence.

III.1 Descriptive Statistics.

Table 1 presents summary statistics of *S**, *S*, and FDI by state. The first column contains the average industry-adjusted index *S**, from 1977 to 1994, as described by equation **(3)**. The

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[10]The state's GSP is in both the numerator and the denominator of **(3)**, so equation **(3)** can be expressed as $S_{st}^* = P_{st}/\hat{P}_{st}$, where $\hat{P}_{st}$ is the summation term in **(2)**.

[11]Support for the inference that relatively high abatement costs indicate relatively stringent regulations can be found in Berman and Bui (1999), which regresses pollution abatement costs at the 4-digit SIC-code level on detailed changes in industry-specific regulations, and finds strong positive associations.

second column contains the unadjusted index, *S*, as described by equation **(1)**.[12] The correlation between the two is about 0.7. A number of states that appear to have relatively high stringency according to the unadjusted index, have much lower ranking after accounting for their industrial compositions. New Jersey, for example, falls from the 20th most stringent state, in column (2) to the 34th in column (1). Other states' apparent stringency improves after controlling for their industries. Florida rises from 25th to 13th. Using the unadjusted measure of standard stringency in column (2), pollution abatement expenditures as a share of gross state product from manufacturing, would give a misleading picture of Florida's and New Jersey's relative environmental compliance costs.

Columns (3) and (4) of Table 1 present the average value of gross property, plant and equipment (PP&E) of foreign-owned affiliates from 1977 to 1994, for all manufacturers and for the chemicals industry, respectively.[13] At the bottom of Table 1 are these same averages for the states with the 5 lowest and 5 highest adjusted pollution abatement indices, S^*, and for the 20 lowest and highest. On average, the five states with the lowest stringency indices have *lower* values of PP&E for foreign-owned affiliates than the five states with the highest indices, and the 20 states with the lowest indices have about the same value of PP&E as the 20 states with the highest stringency. Even looking at SIC code 28, "chemicals and allied products," the five states with the lowest stringency indices have *lower* values of PP&E than the five states with the highest

[12] Because no PACE data were collected in 1987, Table 1 and all subsequent tables omit that year.

[13] We use SIC 28, chemicals and allied products, as an example of a pollution-intensive industry. Of the relatively polluting industries, SIC 28 has the most consistently reported uncensored data in the BEA publications.

indices, and a similar pattern is observable for the 20 lowest and 20 highest states. For many reasons, we would not expect to find a deterrent effect of environmental stringency on FDI as measured by the value of PP&E in these cross-section comparisons. Those states that do not attract a lot of polluting manufacturing probably do not enact stringent regulations -- there is simply less need to worry about industrial pollution in states with less industrial activity, and those states that *do* attract polluting manufacturing may respond by enacting more stringent regulations.

Columns (5) and (6) report similar statistics for employees of foreign-owned affiliates. Here, for all manufacturing and for chemical manufacturing alone, those states with lower environmental compliance costs tend to have more employees. Finally, columns (7) and (8) display the number of planned new foreign-owned plants, from the ITA data. The states with the five lowest stringency indices, and those with the 20 lowest indices, have more annual planned new plant births than the 5 and 20 most stringent, respectively, and this holds true for all manufacturing plants and for the 7 most pollution-intensive 2-digit SIC codes.[14] Again, however, we do not expect these cross-section comparisons to be particularly informative.

The primary advantage of these data over most previous attempts to assess responsiveness to regulatory stringency is their intertemporal variation. Table 2 begins to take advantage of the panel nature of these data by examining *changes* in stringency and FDI. It compares the average stringency and FDI for the first 5 years of the data (1977-1981) to the averages for the last 5 years (1990-1994). The five states whose average stringency fell most during this time period saw their industry-adjusted index of abatement costs fall by 0.597, their average annual value of

[14] The SIC codes included in column (8) are 26 (pulp and paper), 28 (chemicals), 29 (petroleum), 32 (stone clay and glass), 33 (primary metals), and 34 (fabricated metals). These are the industries studied in Co and List (forthcoming).

PP&E grow by $2.5 billion, their average employment in foreign-owned manufacturers grow by 16,698, and their average annual number of new plants grow by 0.32. On the other hand, the five states whose stringency increased the most over the 18 years saw their average index increase by 0.446, their average PP&E grow by only $0.8 billion, their average employment grow by 3,658, and their average number of new plants remain unchanged. While this comparison suggests that states that became more stringent attracted less FDI, the 5 lowest and 5 highest states tend to be the smallest, and much of their variance may be due to noise in the data.

To account for this, the middle two lines of Table 2 examine the 10 states whose relative stringency declined most to the 10 states whose stringency increased most. With the exception of employment in the chemical industry, in every case FDI increased more to those states whose relative environmental stringency declined. The bottom two rows of Table 2 conduct the same exercise for the lowest 20 states and highest 20 states. In general, similar patterns appear, especially for the dirtier industries, though they are muted somewhat by the fact that comparisons among 40 of the 48 continental states necessarily blurs the contrast between states with increasing and decreasing index values.

Table 2 is remarkable, in that it appears to present strong evidence of a deterrent effect of environmental regulations, especially with regard to new plant births in the last two columns. However, the table is based on comparisons that do not control for other observable state characteristics that may have been changing during the same time period. In the next two sections, we control for other such state characteristics, and find that the observed deterrent effect of regulations on FDI largely disappears.

III.2 Estimates using aggregate, continuous data.

To control for other characteristics of states, we estimate variants of

$$FDI_{st} = \beta S^{*}_{st} + \gamma X_{st} + d_{s} + \varepsilon_{st}, \qquad \textbf{(4)}$$

where FDI_{st} is a measure of foreign direct investment in state s during year t, S^{*}_{st} is as defined by equation **(3)**, X_{st} is a set of other state characteristics that may affect investment -- market proximity, taxes, energy costs, land prices, wages, unionization, etc. -- d_s is a set of state-specific time-invariant indicator variables, and ε_{st} is an error term. The state fixed effects, d_s, will account for unobservable state characteristics that would otherwise impart an omitted variable bias.

Table 3 presents the first such estimations. The first column presents means and standard deviations of the regressors. Market proximity is a distance-weighted average of all other states' GSPs. Along with population, this measures the size of the domestic market that may be served by the FDI. Unemployment rates are included to capture labor market characteristics, although of course FDI may affect unemployment simultaneously. Unionization rates measure labor militancy, and may also serve as a regional indicator, since union membership is so much lower in the South. Wages are included as a regressor, though their general positive coefficients may indicate more about productivity than unit labor costs. Total road mileage is included as a measure of public infrastructure, and land prices and energy prices are included to capture factor costs, though they too may be simultaneously determined. Finally, tax effort is an index, calculated as actual tax revenues divided by those that would be collected by a model tax code, as calculated by the Advisory Commission on Intergovernmental Relations (ACIR). Finally, we include a time trend ("year"), and note that we have run most of the specifications using year dummies with almost identical results (not reported).

As a benchmark against which to compare the fixed-effects estimates, columns (2) and (3) contain pooled, OLS regressions of PP&E in the manufacturing sector and the chemical industry, respectively, on the industry-adjusted index of environmental stringency and other covariates, without including state fixed effects (d_s). Controlling for other state characteristics, PP&E appears to be positively correlated with stringency, though the coefficient is insignificant for the chemical industry. However, columns (2) and (3) likely omit state characteristics correlated with both FDI and environmental regulations. Once we include state fixed effects, in columns (4) and (5) the stringency coefficient becomes negative for both manufacturing (−55) and chemicals (−199), and is statistically significant for chemicals.[15]

To interpret the size of these coefficients consider the following. The fixed-effects coefficient in column (5) suggests that a one-unit increase in the stringency index is associated with a decline in chemical industry PP&E by $199 million. The standard deviation of this index within states over time ranges from 0.04 for Wisconsin to 0.56 for Colorado, and averages 0.18. So the coefficient suggests that a one standard deviation increase in the index, for the average state, is associated with a decline in the value of PP&E of foreign-owned chemical manufacturers of $36 million. This amounts to less than 4 percent of the annual average chemical industry PP&E of $1017 million per state.

Table 4 runs some robustness checks on the results described in Table 3. In the first row we run the exact same specifications as in Table 3, but with employment as the dependent variable rather than PP&E. Here the OLS regressions have statistically insignificant stringency

[15]Hausman tests of random effects versus fixed effects models consistently reject the assumption that the error terms are uncorrelated with the regressors.

coefficients, while again the fixed-effects regression for the chemical industry has the only negative and statistically significant stringency coefficient (−991). If we take literally the coefficient for the chemical industry, it suggests that a one-standard-deviation increase in a state's environmental stringency index (+0.18) is associated with 178 fewer jobs in that industry, a fall of about 2.3 percent relative to the average of 7692 employees in foreign-owned chemical plants per state.

In the second row of Table 4 we estimated the same set of regressions from Table 3 using the unadjusted index *S*. This is the index that has often been used by the literature without controlling for states' industrial compositions. The fixed-effects stringency coefficient for the chemical industry (−4597) is less statistically significant, which we would expect if the industrial composition of some states is making them appear less stringent than they actually are. In addition, the point estimate is even smaller than in the case of the industry-adjusted index. The average within-state standard deviation of the unadjusted index is 0.0034. A one standard deviation in the unadjusted index, for the average state, is therefore associated with a decline in the value of PP&E of foreign-owned chemical manufacturers of $15 million, about 1.5 percent of average annual chemical industry PP&E. The third row of Table 4 contains the same specification, using employment in foreign-owned manufacturing facilities as the dependent variable, and has a similar pattern of coefficients. The unadjusted index of compliance costs again yields coefficients that are biased upwards.

To address concerns that year-to-year noise in the data mask long-run trends, in rows (4) and (5) of Table 4 we estimate specifications based on averages of the data over three time periods: 1977-81, 1982-86, and 1988-94. The 48 states, and three time periods, comprise 144 observations. The pattern of coefficients is largely similar to those using the annual data. The pooled specifications in columns (1) and (2) yield positive coefficients that are insignificant for the industry-adjusted index (*S**), significant for the unadjusted index (*S*). The fixed-effects specifications (with only three time periods) in columns (3) and (4) yield mostly negative and insignificant coefficients.

So far, the evidence presented has all been based on a static model of investment in which annual measures of FDI are regressed on concurrent state characteristics. However, one might object that investment is by nature a dynamic process. FDI may, for example, be a function of expected future state characteristics. In addition, FDI to existing facilities will be a function of past investments to those facilities. In either case, the usual orthogonality conditions may not hold across time. To explore this issue in a dynamic context, suppose that a reduced form relationship for FDI can be characterized by the following equation:[16]

$$FDI_{st} = \rho FDI_{s,t-1} + \beta S^{*}_{st} + \gamma X_{st} + d_s + \varepsilon_{st} \qquad \textbf{(5)}$$

Equation **(5)** states that *FDI* is a function of current state characteristics and lagged values of *FDI*. Both FDI_{st} and $FDI_{s,t-1}$ are functions of d_s, a part of the unobserved error term, and therefore OLS fixed-effects estimates of **(5)** will be biased because $FDI_{s,t-1}$, a regressor, is correlated with the error term (Amemiya 1985).

Arellano and Bond (1991) suggest a GMM estimation of **(5)** that uses lagged values of $FDI_{s,t-1}$ as instruments for $FDI_{s,t-1}$, and first differences to eliminate the fixed state effects d_s:

$$\Delta FDI_{st} = \rho \Delta FDI_{s,t-1} + \beta \Delta S*_{st} + \gamma \Delta X_{st} + \Delta \varepsilon_{st} \qquad \textbf{(6)}$$

[16]This discussion is based on Baltagi (1995) and Arellano and Bond (1991).

where Δ symbolizes first differences. Since $FDI_{s,t-2}$ is correlated with $\Delta FDI_{s,t-1}$, but not correlated with ΔFDI_{st}, it is a valid instrument. In fact, all past values $FDI_{s,t-3}$, $FDI_{s,t-4}$, and so on, as well as values of the exogenous variables S^* and X, are valid instruments for $\Delta FDI_{s,t-1}$.

Row (6) of Table 4 presents the coefficient β from GMM estimates of **(6)** using the Arellano and Bond estimator as programmed in Doornik *et al.* (1999).[17] When equation **(6)** is estimated using all manufacturing FDI, in column (3), the coefficient (2.4) is tiny and statistically insignificant, though still positive. Turning to the chemical industry, in column (4), the coefficient (−338) is negative and statistically significant, and 70 percent larger than the fixed-effects estimate (-199) from Table 3. Even this larger estimate, however, implies that a one-standard-deviation increase in compliance costs is associated with decline in FDI of less than 6 percent. Compared with the pooled cross-section analyses in columns (1) and (2) of Table 3, this provides more evidence that the positive coefficients found in cross-section analyses are spurious, and are based on unobserved characteristics correlated with both environmental regulations and economic activity.[18]

In sum, using continuous data on investment and employment by foreign-owned manufacturers in U.S. states, we do not find broad evidence that stringent environmental regulations are reducing FDI overall, though they may be affecting FDI by particularly pollution-

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[17]Doornik, Bond, and Arellano's GMM estimation is written for the computer package Ox, and may be downloaded from http://www.nuff.ox.ac.uk/Users/Doornik/. See Doornik (1998) and Doornik *et al.* (1999).

[18]It is also notable that the parameter ρ on lagged FDI is precisely estimated at about 0.9 in these specifications. This means that states hosting relatively high levels of FDI in any given year experience slower growth of FDI in subsequent years, implying that, *ceteris paribus*, the levels of FDI are converging across U.S. states.

intensive industries. However, even when we estimate statistically significant coefficients, their implied magnitudes are generally small economically. These results are obtained from an analysis that has addressed three important problems that pervade much of the previous literature: it examines inflows of FDI to U.S. states; it uses a panel of data to account for unobserved heterogeneity among states; and it uses a quantitative measure of stringency based on compliance costs, adjusting for states' industrial compositions.

Despite these strengths, the measure of FDI used thus far is not without a few weaknesses. One important problem is that changes in observed FDI can result from new plants being constructed, old plants being closed, or from expansions and contractions of existing plants. Each of these four types of changes may respond quite differently to changes in environmental regulations, depending on how the regulations are written. Many state environmental regulations consist of "new source performance standards" that are more stringent for new plants than for existing plants. These standards effectively raise barriers to entry that protect existing older plants. Using aggregate data may conceal effects that work in opposite directions. A second problem with the BEA data is that they include FDI in the form of mergers and acquisitions. If future environmental compliance costs are capitalized into the prices paid for acquisitions, then cost differences among states will be exactly offset by price differences and will in theory have no effect on FDI. Consequently, in order to isolate the effects of regulations on the location of FDI, without the offsetting effect of grandfather regulations or cost capitalization for existing investment, in the next section we use establishment-level data to focus on new plants only.

III.3 Estimates using establishment-level count data.

To examine FDI in new plants only, we turn to the International Trade Administration (ITA) data. Though the data include acquisitions, mergers, joint ventures, real estate transactions, equity increases, plant expansions and new plants, we focus only on the new plants. Because the ITA data do not come from a mandatory survey, they may miss some foreign investment. However, the ITA's claim that its data cover "the vast majority of significant foreign direct investment transactions" is confirmed by BEA officials.[19]

We begin by estimating the effect of environmental regulations on FDI using the basic Poisson specification:

$$prob(n_{st}) = \frac{e^{-\lambda_{st}} \lambda_{st}^{n_{st}}}{n_{st}!} \qquad \textbf{(7)}$$

where n_{st} is the number of new plants in state s in year t, and λ is the Poisson mean and variance. We make the standard assumption that log $\lambda_{st} = X_{st}\beta + \varepsilon_{st}$, where X_{st} are state characteristics and β is a vector of parameters to be estimated via maximum likelihood.[20]

Table 5 contains estimates of β. When the data are pooled, the coefficient on the industry-adjusted index of environmental regulatory stringency (−0.049) is negative but not significant. Furthermore, the coefficient suggests that a one-standard-deviation increase in the index would be

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[19]Personal correspondence. The ITA data come from newspapers, magazines, and business and trade journals, as well as from public files of Federal regulatory agencies such as the Securities and Exchange Commission, the Federal Trade Commission, and the Federal Reserve Board.

[20]Gourieroux, Monfort and Trognon (1984) have shown that the parameters β are consistently estimated even if the Poisson distributional assumption does not hold, as long as the conditional mean, logλ_{st}=$X_{st}\beta$, is correctly specified.

associated with only a 0.9 percent decline in the annual number of new plants. Turning to the polluting industries in column (2), the stringency coefficient is *positive*, though again statistically insignificant.

However, the pooled specifications in columns (1) and (2) make no use of the panel of data, and are almost certainly misspecified, since the error terms ε_{st} are likely to be correlated within states. Therefore, in columns (3) and (4) we estimate a fixed-effects Poisson regression, based on Hausman, Hall and Griliches (1984). For all manufacturing industries, in column (3), the stringency coefficient remains insignificant, though for the sample of polluting industries the stringency coefficient is *positive* and significant. This finding is puzzling, especially since it appears to contradict the simple differences of means in Table 2 and the findings based on the BEA data on PP&E and employment in Table 3.

In Table 6 we present some alternative count-data models. Because the Poisson regressions impose the strong assumption that the mean equals the variance (λ), in the first row we estimate the same models using a negative binomial specification. Here the fixed effects models are not fixed effects in the ordinary sense of the term, but rather refer to the distribution of the dispersion parameter, not to the $X_{st}\beta$ term. The results largely parallel those of the more restrictive Poisson regression. The stringency coefficient is negative, small, and statistically insignificant for the pooled data, and positive, small and insignificant for the fixed-effects models, though curiously it is close to statistical significance for the polluting industries.[21]

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[21]We have also used the ITA data to estimate a fixed-effects multinomial logit model (Chamberlain, 1980). The pattern of environmental stringency coefficients (not reported here) for all manufacturers and dirty industries, and for the adjusted and unadjusted indices, parallels that for the Poisson regressions.

In the second and third rows of Table 6 we estimate the count data models using the *unadjusted* index of stringency. In most cases the stringency coefficients are positive, much larger, and more statistically significant than those for the *industry-adjusted* index. This result confirms our suspicions about the unadjusted index used by others: it is higher for states that have a lot of investment in dirty industries, and is therefore more of an indicator of polluting industrial compositions than stringent regulations.

In columns (4) and (5) of Table 6 we present estimates based on the multi-year averages of the data, to address concerns that year-to-year noise masks substantive changes. The pooled specifications in columns (1) and (2) yield negative, statistically significant coefficients, and these shrink and become insignificant once fixed effects are added in columns (3) and (4). Again, the pattern is exaggerated in row (5), which uses the unadjusted index (*S*) and is therefore biased in favor of finding a positive association between FDI and compliance costs.

Finally, one might be concerned that the Poisson regressions are biased because so many of the states had zero plant births in any given year. Of the 768 state-year observations, 412 experienced zero plant births during the 16 years, and 519 experience zero births in the polluting industries. Consequently, in rows (6) and (7) of Table 6 we estimate a "zero-inflated Poisson" (also called a "hurdle model") version of the basic pooled specifications (Greene, 1997). These assume that the number of new plants in a state, n_{st}, is governed by the following process:

$$\begin{aligned} prob(n_{st}=0) &= e^{-\theta} \\ prob(n_{st}=n) &= \frac{(1-e^{-\theta})e^{-\lambda}\lambda^{n}}{n!(1-e^{-\lambda})} \end{aligned} \qquad \textbf{(8)}$$

The specifications at the bottom of Table 6 use a logit model to estimate the top equation, with state populations, market proximity, unionization rates, and road mileage as regressors. The results are similar to those from the basic Poisson regression in Table 5 and in column (2) of Table 6. The stringency index is insignificant in the pooled data for all manufacturing and for the polluting industries. We conclude from these pooled models that the many states with zero new plant births in some years are not driving the results. As with almost every other specification, using the unadjusted measure of compliance costs, *S*, in row (7) upwardly biases the compliance cost coefficient.

IV. Discussion and Implications.

Before drawing conclusions based on these results, we must acknowledge several important caveats. First, our industry-adjusted index of environmental stringency, *S**, controls for states' industrial compositions at the level of 2-digit SIC codes. While this surely accounts for a lot of the differences among states, there is equally certain to be heterogeneity among states *within* 2-digit classifications. For example, industry code 26, pulp and paper, contains paper mills, which are among the most pollution-intensive manufacturers, along with envelope assemblers, which emit very little pollution. To the extent that some states contain relatively more pulp mills and others merely assemble envelopes, high abatement costs in the former will not necessarily reflect more stringent environmental regulations. Consequently, one explanation for this paper's failure to measure a large, robust deterrent effect of environmental regulations on investment is that the 2-digit industry adjustment may still mask considerable heterogeneity, and that states that find themselves attracting relatively polluting industries -- *within* any given 2-digit SIC code -- may respond by enacting strict regulations. While this concern certainly merits inquiry, such an analysis is beyond the scope of this paper. Furthermore, by controlling for state industrial compositions at the 2-digit level we have both pointed out the importance of this measurement issue, and have controlled for a significant fraction of heterogeneity in state industrial compositions.

A second caveat involves the efforts that states make to attract and retain certain industries. These efforts are largely unmeasured in the current estimations. However, one can easily imagine that changes in state efforts to promote investment in particularly polluting industries may be correlated with environmental regulations affecting those industries. It may be that states enacting stringent environmental regulations enact compensatory tax breaks or infrastructure subsidies. Or, it may be that states enacting weak pollution regulations are also inclined to pass generous investment subsidies. Under the former circumstances, we are likely to have underestimated the deterrent effect of regulations, absent the development incentives. Under the latter scenario, we may be overstating the effect of environmental regulations by falsely attributing some of the effects of unobserved development incentives to correlated observed low environmental costs. Again, analysis of the political economies of state pollution regulations lies outside our agenda for this paper.

Third, our industry-adjusted index makes no attempt to control for the relative *age* of different states' manufacturers. This is important because many state environmental standards are more strict for new sources of pollution than for existing sources. Consequently, states such as Florida, that have relatively new manufacturing bases, have relatively high compliance costs, even after controlling for their industrial compositions. Conversely, states such as Connecticut that have relatively old manufacturers will experience lower compliance costs. There is, therefore, a potential positive correlation between the amount of new investment and our industry-adjusted index of regulatory stringency. Evidence suggests, however, that such a correlation may be insignificant (Levinson, 1996).

In studying the effect of differences in environmental compliance costs on the location of inward FDI to U.S. states, our approach has two distinctive features. First, our measure of stringency controls for the industrial composition of states. The results indicate that this is important: both the least-squares and the Poisson regressions yield stronger and more positive association between environmental regulations and FDI with the unadjusted, compared to the adjusted abatement cost index. This suggests that a high unadjusted abatement cost index primarily reflects a high share of industrial activity in polluting industries. Therefore, results from studies that do not take this composition effect into account allow only very limited inferences.

Second, our panel approach controls for unobserved heterogeneity through the inclusion of fixed effects. It is not clear *a priori* which way the omitted-variable bias, if any, goes, but with the continuous PP&E and employment data for all investment in Tables 3 and 4, the coefficients on the environmental variables *fall* when state fixed effects are added to the model. For the count data on new investment in Tables 5 and 6, however, the environmental coefficients *rise* with the inclusion of the state fixed effects.

In addition to the literature on the effects of environmental regulation on FDI, this paper is related to work examining the effect of interstate tax differentials on new firm location in the United States and on the effect of tax differences across the U.S. states on FDI inflows. Hines (1996) studies the effects of corporate income tax differences in U.S. states on inward FDI by comparing investment from the U.K. and Japan with that from a set of other OECD countries, where differences in their respective foreign tax credit systems make the former two countries less

"US tax sensitive" than the other countries. Based on this identifying assumption, which also goes some way to addressing the omitted variable problem, Hines estimates much stronger effects from income taxation on FDI location than we have done here.[22]

Bartik (1985) and Papke (1991) are two examples of recent work on the effect of income taxation on *domestic* investment in the U.S. states. The former estimates tax elasticities of -20 to -30 percent, but his specification based on a cross-section of data around 1972 does not fully address issues of unobserved heterogeneity.[23] Papke (1991) also employs a fixed-effects Poisson specification with panel data, estimating elasticities of firm births with respect to marginal effective income tax rate increases between -160 and -1570 percent for certain industries. Her estimates are also far larger than what we have estimated for environmental regulations. Thus, while changes in income taxation across U.S. states might trigger large measurable responses in the location of firms' investments, the same cannot be said about the effect of changes in environmental compliance costs.

Finally, while the motivation for this research is to draw inferences about the sensitivity of FDI to *international* differences in environmental stringency, we recognize that the stringency of environmental legislation differs much more across countries than across U.S. states. However, the variation in other characteristics such as factor costs, market access, transportation costs, and exchange rate risks also varies more across countries than across states. Thus, our analysis does

[22]Specifically, for every one percent increase in taxation, the share of FDI owned by "US tax-sensitive" countries drops by ten percent, and the share of these countries in the number of total foreign affiliates drops by three percent.

[23]Moreover, even the nested logit specification that Bartik employs does not fully address the problems associated with the 'Independence of Irrelevant Alternatives' assumption.

not necessarily underestimate the sensitivity of FDI location with respect to environmental legislation at the international level.

In sum, despite three important improvements over the existing literature, the results presented here largely confirm the perception among economists who have looked for statistical evidence of the deterrent effect of environmental regulations on economic activity. Such effects, if they exist at all, appear to be small and confined to a few polluting industries. We do feel, however, that we have assembled the best possible data set with which to ask this important question, and that we have eliminated several potential explanations for the lack of statistical evidence. By looking at FDI inflows to U.S. states we examine comparable jurisdictions with comparable environmental compliance cost data. By adjusting for those states' industrial compositions, we eliminate bias caused by the uneven distribution of industries among states. And by examining FDI and environmental stringency using a 17-year panel, we control for potential unobserved heterogeneity among states that may be correlated with both the amount of FDI and the stringency of state regulations.

References

Amemiya, Takeshi. 1985. Advanced Econometrics. Cambridge, MA. Harvard University Press.

Arellano, Manuel, and Bond, Stephen R. 1991. "Some test of specification for panel data: Monte Carlo evidence and an application to employment equations" Review of Economic Studies 58, 277-297.

Baldwin, Robert E. 1995. "An Economic Evaluation of the Uruguay Round Agreements," S. Arndt and C. Milner (eds.), The World Economy: Global Trade Policy 1995, Blackwell Publications, Oxford, pp. 153-172.

Baltagi, Badi H. 1995. Econometric Analysis of Panel Data. New York: John Wiley & Sons.

Bartik, Timothy. 1988. "The Effects of Environmental Regulation on Business Location in the United States," Growth and Change Summer 22-44.

Bartik, Timothy. 1985. "Business Location Decisions in the United States: Estimates of the Effects of Unionization, Taxes, and other Characteristics of States," Journal of Business & Economic Statistics Vol.3, 1: 14-22.

Becker, Randy and Vernon Henderson. 1977. "Effects of Air Quality Regulation on Decisions of Firms in Polluting Industries," NBER working paper #6160, September.

Berman, Eli and Linda Bui. 1999. "Environmental Regulation and Productivity: Evidence from Oil Refineries," mimeo, Boston University.

Bhagwati, Jagdish N., and T.N. Srinivasan. 1995. "Trade and the Environment: Does Environmental Diversity Detract from the Case for Free Trade?" Economic Growth Center Working Paper No.721, Yale University.

Blundell, Richard, Stephen Bond, and Costas Meghir. 1996. "Econometric Models of Company Investment" in Laszlo Matyas and Patrick Sevestre, eds. The Econometrics of Panel Data. Boston: Kluwer Academic Publishers.

Brainard, Lael. 1993. "A Simple Theory of Multinational Corporations and Trade with a Trade-off between Proximity and Concentration," NBER WP # 4269, Cambridge, MA.

Bureau of Economic Analysis (BEA). 1999. Foreign Direct Investment in the United States, annual series, various years, U.S. Department of Commerce, Washington, D.C.

Bureau of Economic Analysis (BEA). 1997. "U.S. Intra-firm Trade in Goods," Survey of Current Business, Bureau of Economic Analysis, February.

Caves, Richard E. 1996. Multinational Enterprise and Economic Analysis, 2nd edition, Cambridge University Press, 1996.

Chamberlain, Gary. 1980. "Analysis of covariance with qualitative data," Review of Economic Studies 47, 225-238.

Co, Catherine and John List. 1998. "The Effects of Environmental Regulations on Foreign Direct Investment," Journal of Environmental Economics and Management, forthcoming.

Copeland, Brian R. 1990. "Strategic Interaction Among Nations: Negotiable and Non-negotiable Trade Barriers," Canadian Journal of Economics, 23: 64-108.

Crandall, Robert W. 1993. Manufacturing on the Move. Washington, DC: Brookings Institution.

Cropper, Maureen and Wallace Oates. 1992. "Environmental Economics: A Survey," Journal of Economic Literature Vol. 30, pp.675-740.

Doornik, Jurgen A. 1998. Object-Oriented Matrix Programming using Ox 2.0. London: Timberlake Consultants Ltd and Oxford: www.nuff.ox.ac.uk/Users/Doornik.

Doornik, Jurgen A., Manuel Arellano, and Stephen Bond. 1999. "Panel Data estimation using DPD for Ox," mimeo.

Dunning, John H. 1981. International Production and the Multinational Enterprise. London: George Allen and Unwin.

Ederington, Josh, 1998, "International Trade and Domestic Policies," mimeo, University of Miami.

Friedman, Joseph, Daniel A. Gerlowski and Jonathan Silberman. 1992. "What Attracts Foreign Multinational Corporations? Evidence From Branch Plant Location in the United States," Journal of Regional Science Vol. 32, No. 4, pp.403-418.

Gourieroux, Christian, Alain Monfort and Alain Trognon. 1984. "Pseudo maximum likelihood methods: Theory," Econometrica 52: 681-700.

Gray, Wayne B. 1997. "Manufacturing plant location: Does state pollution regulation matter?" NBER Working Paper #5880.

Greene, William H. 1997. Econometric Analysis. Third Edition. Prentice Hall: New Jersey.

Greenstone, Michael. 1998. "The Impacts of Environmental Regulations on Industrial Activity: Evidence from the 1970 and 1977 Clean Air Act Amendments and the Census of Manufactures," mimeo, University of California at Berkeley.

Hausman, Jerry, Bronwyn H. Hall, and Zvi Griliches. 1984. "Econometric Models for Count Data with an Application to the Patents-R&D Relationship," Econometrica 52(4) 909-938.

Head, Keith, John Ries, and Deborah Swenson. 1995. "Agglomeration benefits and location choice: Evidence from Japanese manufacturing investments in the United States." Journal of International Economics 38(3-4) May, 223-247.

Henderson, J. Vernon. 1996. "Effects of Air Quality Regulation," American Economic Review 86(4), September, 789-813.

Hines, James R, Jr. 1996. "Altered States: Taxes and the Location of Foreign Direct Investment in America," American Economic Review 86(5): 1076-1094.

Horstmann, Ignatius, and James Markusen. 1992. "Endogenous Market Structures in International Trade," Journal of International Economics 32: 109-129.

International Trade Administration (ITA). Foreign Direct Investment in the United States (various years). United States Department of Commerce. Office of Trade and Economic Analysis.

Jaffe, Adam B., Steven R. Peterson, Paul R. Portney, and Robert N. Stavins. 1995. "Environmental Regulations and the Competitiveness of U.S. Manufacturing: What does the Evidence Tell Us?" Journal of Economic Literature 33(1):132-63.

Kahn, Matthew. 1997. "Particulate Pollution Trends in the U.S." Regional Science and Urban Economics, 27: 87-107.

Keller, Wolfgang. 1998. "Multinational enterprises, international trade, and technology diffusion" , mimeo, University of Texas-Austin, August.

Kolstad, Charles, and Y. Xing. 1997. "Do Lax Environmental Regulations Attract Foreign Investment?" UC Santa Barbara Working Paper #6-95R, February.

Leonard, H. Jeffrey. 1988. Pollution and the Struggle for the World Product, Cambridge: Cambridge University Press.

Levinson, Arik. 1996. "Environmental Regulations and Manufacturers' Location Choices: Evidence from the Census of Manufactures," Journal of Public Economics, 61(1).

Levinson, Arik. 1999. "An Industry-Adjusted Index of State Environmental Standard Stringency," mimeo, University of Wisconsin Economics Department.

List, John A. 1999. "U.S. County-Level Determinants of Inbound FDI: Evidence from a Two-Step Modified Count Data Model," International Journal of Industrial Organization, forthcoming.

List, John A., and Mitch Kunce. 1999. "Environmental Protection and Economic Growth: What do the Residuals tell us?" Land Economics, forthcoming.

Lipsey, Robert, Magnus Blomstrom, and Erich Ramstetter. 1998. , "Internationalized Production in the World Economy,"in Robert Baldwin et al. (eds.), Geography and Ownership as Bases for Economic Accounting, The University of Chicago Press for the NBER, pp.83-135.

Low, Patrick, and Alexander Yeats. 1992. "Do 'Dirty' Industries Migrate?" In Patrick Low, ed., International Trade and the Environment. Washington, DC: World Bank.

Low, Patrick, ed. 1992. International Trade and the Environment, Washington, DC: The World Bank.

Mataloni, Ray. 1999. "A Guide to BEA Statistics on U.S. Multinational Companies," in International Direct Investment Studies by the Bureau of Economic Analysis, U.S. Department of Commerce, pp.198-215; Table 3.

Papke, Leslie E. 1991. "Interstate business tax differentials and new firm location," Journal of Public Economics 45: 47-68.

Rauscher, Michael. 1997. International Trade, Factor Movements, and the Environment. Oxford and New York: Oxford University Press, Clarendon Press.

Table 1: Summary Statistics
Averages 1977-1994

State	Compliance cost index S* (1)	Unadjusted index S (2)	Property, Plant & Equipment of foreign-owned affiliates ($millions): Manufacturing (3)	Chemicals (4)	Employees of foreign-owned affiliates: Manufacturing (5)	Chemicals (6)	Annual number of new foreign-owned plants: Manufacturing (7)	Polluting industries[e] (8)
Alabama	1.19	0.0219	2876	803	22747	4502	0.94	0.65
Arizona	1.39	0.0148	1430	206	11927	2588	0.35	0.18
Arkansas	1.17	0.0168	826	131	14152	2034	0.18	0.12
California	0.90	0.0121	10397	2026	150365	33285	5.24	2.06
Colorado	1.01	0.0113	926	320	13612	3085	0.35	0.12
Connecticut	0.67	0.0079	1565	335	29448	4825	0.88	0.35
Delaware	1.30	0.0344	2786	2724	25304	32300	0.65	0.59
Florida	1.21	0.0138	2940	749	43306	6878	0.76	0.24
Georgia	0.91	0.0127	4729	861	54975	8947	3.82	1.00
Idaho	1.66	0.0181	210	24	3138	434	0.00	0.00
Illinois	0.91	0.0132	6389	1331	86496	14230	2.29	1.06
Indiana	1.14	0.0196	5088	765	53453	8609	2.18	1.41
Iowa	0.96	0.0106	1245	262	15482	3406	0.53	0.24
Kansas	0.76	0.0115	705	182	10828	2420	0.24	0.06
Kentucky	0.99	0.0146	2923	561	25185	4289	1.76	1.00
Louisiana	1.51	0.0538	5094	2835	18421	6974	0.47	0.41
Maine	1.55	0.0237	1093	42	8713	449	0.12	0.06
Maryland	1.17	0.0185	1799	408	26491	5484	1.00	0.24
Massachusetts	0.67	0.0067	2126	506	39880	8212	1.00	0.35
Michigan	1.01	0.0121	4129	631	55779	7827	2.12	1.18
Minnesota	0.66	0.0092	1720	168	24294	3522	0.18	0.12
Mississippi	1.47	0.0213	990	518	10585	1651	0.29	0.18
Missouri	0.79	0.0104	2404	664	27731	7312	0.71	0.53
Montana	1.49	0.0341	528	566	1496	554	0.00	0.00

(continued)

Table 1 (continued)

State	Compliance cost index S* (1)	Unadjusted index S (2)	Property, Plant & Equipment of foreign-owned affiliates ($millions)		Employees of foreign-owned affiliates		Annual number of new foreign-owned plants	
			Manufacturing (3)	Chemicals (4)	Manufacturing (5)	Chemicals (6)	Manufacturing (7)	Polluting industries[e] (8)
Nebraska	0.83	0.0088	257	72	5226	1502	0.06	0.00
Nevada	0.63	0.0072	270	38	2944	650	0.29	0.18
New Hampshire	0.75	0.0072	492	35	9999	505	0.00	0.00
New Jersey	0.82	0.0158	6972	3810	88583	43431	2.47	1.65
New Mexico	1.64	0.0306	679		2701	607	0.12	0.00
New York	0.77	0.0087	5055	1084	91944	17760	4.59	1.65
N. Carolina	0.82	0.0088	6485	2467	76700	22005	4.12	2.12
N. Dakota	0.77	0.0105	189		1448	417	0.00	0.00
Ohio	0.82	0.0139	6177	1044	83174	11386	2.82	2.29
Oklahoma	0.58	0.0103	1614	1296	13929	5106	0.24	0.18
Oregon	1.22	0.0139	871	88	9559	1491	1.00	0.47
Pennsylvania	0.91	0.0169	5891	1450	92059	17095	1.94	1.18
Rhode Island	0.72	0.0075	506	151	7577	1401	0.18	0.00
S. Carolina	0.99	0.0160	4913	2056	44540	13793	1.71	0.82
S. Dakota	0.68	0.0056	62	4	1601	113	0.00	0.00
Tennessee	1.10	0.0165	4554	1480	52981	12125	2.41	0.94
Texas	1.39	0.0311	14632	7970	89008	25756	3.82	2.88
Utah	0.93	0.0164	480	120	7117	1077	0.12	0.00
Vermont	0.66	0.0065	215	9	2729	131	0.35	0.18
Virginia	0.96	0.0118	3295	1637	36171	11360	2.53	1.00
Washington	1.37	0.0196	2197	153	18107	2874	0.88	0.18
W. Virginia	1.58	0.0433	3024	2229	16123	8772	0.18	0.18
Wisconsin	0.89	0.0110	2161	154	38627	4142	0.47	0.24
Wyoming	0.72	0.0259	838	992	1225	1005	0.00	0.00
Avg. for lowest 5[a]	0.64	0.0082	1077	293	14669	2931	0.39	0.20
Avg. for highest 5[b]	1.59	0.0339	2020	1314	9819	3261	0.18	0.13
Avg. for lowest 20[c]	0.75	0.0103	2525	803	35413	8987	1.19	0.60
Avg. for highest 20[d]	1.33	0.0235	2841	1229	24874	5609	0.89	0.50

Omits AK, HI, and 1987. Columns (4) and (6) omit 1992-94, and columns (7) and (8) omit 1989.
[a]OK, NV, MN, CT, MA.
[b]NM, ID, WV, ME, MT.
[c]Add to (a) VT, SD, RI, NH, KS, NY, ND, MO, NJ, NC, WY, OH, NE, WI, GA.
[d]Add to (b) LA, MS, TX, AZ, WA, DE, OR, FL, AL, MD, AR, IN, TN, MI, CO.
[e]SIC codes 26, 28, 29, 32, 33, 34, 37. (See footnote 15.)

Table 2
Changes in Average Pollution Abatement Costs and FDI
(1977-1981) to (1990-1994)

	Change in industry-adjusted index of abatement costs (S^*) (1)	Property, plant & equipment change		Employment change		Total new plants	
		total manufacturing (2)	chemical industry (3)	total manufacturing (4)	chemical industry (5)	total manufacturing (6)	polluting industries[g] (7)
Lowest 5[a]	−0.597	2,495	1,311	16,698	1,306	0.32	0.00
Highest 5[b]	0.446	801	209	3,658	451	0.00	−0.04
Lowest 10[c]	−0.370	3660	982	20,949	2206	0.58	0.28
Highest 10[d]	0.310	3007	720	19,567	2972	−0.44	−0.20
Lowest 20[e]	−0.230	4,508	1,757	26,183	5,796	0.43	0.13
Highest 20[f]	0.190	5,282	1,551	31,577	4,385	−0.13	0.00

[a]AZ, NM, ID, DE, FL.
[b]WY, ND, RI, CO, SD.
[c]Add to (a) IN, AL, IA, WA, OK.
[d]Add to (b) ME, CT, MA, IL, GA.
[e]Add to (c) NJ, WV, MS, OR, MI, PA, MT, MD, VA, NC.
[f]Add to (d) MN, CA, TX, SC, UT, OH, WI, NY, NH, KY.
[g]SIC codes 26, 28, 29, 32, 33, 34, 37. (See footnote 15.)

Table 3
Value of Gross Property, Plant and Equipment (PP&E)
As a Function of Abatement Costs
1977 – 1994

	Mean (std. dev.) (1)	Pooled Manufacturing (2)	Pooled Chemicals[a] (3)	State Fixed Effects Manufacturing (4)	State Fixed Effects Chemicals[a] (5)
Industry-adjusted index of abatement costs (*S**)		500* (237)	267 (186)	–55 (219)	–199* (161)
Market proximity	6631 (8220)	0.207* (0.019)	0.098* (0.015)	0.376* (0.022)	0.079* (0.016)
Population (1000s)	4940 (5134)	0.175* (0.033)	–0.016 (0.023)	1.16* (0.10)	0.389* (0.071)
Unemployment rate	6.61 (2.09)	122* (43)	86.0* (29.1)	–24.9 (36.6)	33.3 (24.8)
Unionization rate	16.6 (6.7)	–108* (20)	–84.6* (13.9)	–92.7* (41.6)	–86.0* (29.1)
Wages	9.10 (2.24)	179* (87)	32.9 (66.7)	310* (151)	211† (114
Road mileage (1000s)	80.5 (48.4)	12.3* (2.6)	10.8* (1.8)	55.5* (11.7)	53.7* (8.2)
Land prices (per acre)	887 (775)	0.52* (0.12)	0.62* (0.10)	–0.31† (0.17)	0.16 (0.13)
Energy prices	5.51 (1.70)	–288* (56)	–144* (41)	–259* (57)	–139* (40)
Tax effort	96.1 (16.1)	–31.0* (5.9)	–11.4* (4.1)	47.6 (8.7)	25.6* (6.2)
Year		166* (41)	32.4 (33.4)	–4.0 (48.5)	–46.0 (38.4)
Constant		–11602* (3072)	–1525 (2516)	–13476* (3397)	–4530† (2734)
no. observations	816	811	563	811	563
no. censored		5	109	5	109
R^2		0.70	0.47	0.77	0.52

Standard errors in parentheses.
1987 is dropped because no PACE data were collected that year.
* Statistically significant at 5 percent.
† Statistically significant at 10 percent.
[a] The chemical industry investment data is only for 1977-1991.

Table 4
Alternative measures of regulatory stringency:
Gross Property, Plant and Equipment and Employment
as a function of abatement costs
1977 – 1994

		Pooled		State Fixed Effects	
	Coefficients on Index of Abatement Costs	Manufacturing (1)	Chemicals[a] (2)	Manufacturing (3)	Chemicals[a] (4)
(1)	Adjusted index (*S**), with employment as the dependent variable.	–2288 (1586)	–825 (755)	549 (1301)	–991* (492)
(2)	Unadjusted Index (*S*), with PP&E as the dependent variable.	66589* (8803)	72215* (6608)	13651 (11570)	–4597 (10105)
(3)	Unadjusted index (*S*), with employment as the dependent variable.	46548 (61563)	106909* (29767)	107379 (68320)	–27580 (29301)
(4)	Five-year averages, with PP&E and adjusted index (*S**).[b]	483 (614)	440 (462)	–187 (817)	–669 (596)
(5)	Five-year averages, with PP&E and unadjusted index (*S*).[b]	72750* (21889)	85282* (14326)	17658 (43455)	–33695 (36060)
(6)	Dynamic Panel Model (GMM), with PP&E and adjusted index (*S**).	--	--	2.4 (92.6)	-338* (100)

Standard errors in parentheses.
* Statistically significant at 5 percent.
† Statistically significant at 10 percent.
[a] The chemical industry investment data is only for 1977-1991.
[b] Rows (4) and (5) average all dependent and independent variables for 1977-1981, 1982-1986, and 1988-1994, and treats each period as one observation. There are thus 48 states and three periods, for 144 total observations. The last period, 1988-1994, takes a seven-year average for total manufacturing in columns (1) and (3), and a 4-year average for the chemical industry, in columns (2) and (4).

Table 5
Count Data Models
Of New Foreign-owned Plants
As a Function of Abatement Costs
1977 – 1994

Regressors	Pooled Poisson: All Manufacturing (1)	Pooled Poisson: Polluting industries (2)	Fixed-Effects Poisson: All manufacturing[a] (3)	Fixed-Effects Poisson: Polluting industries[b,c] (4)
Industry-adjusted index of abatement costs	–0.049 (0.121) [0.952]	0.031 (0.168) [1.032]	0.379 (0.263) [1.460]	0.776* (0.367) [2.173]
Market proximity (millions)	8.14 (5.31)	16.2* (7.8)	–76.8* (13.8)	–84.6* (21.7)
Population (millions)	0.077* (0.008)	0.050* (0.012)	0.178* (0.046)	0.128† (0.072)
Unemployment rate	–0.069* (0.020)	–0.039 (0.028)	–0.194* (0.030)	–0.184* (0.042)
Unionization rate	–0.019† (0.010)	–0.012 (0.014)	–0.010 (0.034)	–0.079† (0.047)
Wages	–0.011 (0.042)	0.080 (0.058)	0.274* (0.133)	0.120 (0.183)
Road mileage (millions)	2.93* (0.92)	5.35* (1.25)	–2.43 (8.40)	6.78 (11.5)
Land prices ($1000 per acre)	0.095† (0.056)	0.136† (0.075)	–0.592* (0.164)	–0.345 (0.225)
Energy prices	0.055* (0.026)	0.040 (0.037)	0.145* (0.048)	0.204* (0.070)
Tax effort	0.0102* (0.0025)	0.0045 (0.0036)	0.0155* (0.0069)	0.024* (0.010)
Year	–0.015 (0.020)	–0.0484† (0.0277)	–0.0551 (0.0421)	–0.043 (0.059)
Constant	0.159* (1.47)	1.58 (2.10)	na	na
n	768	768	672	608
Pseudo R^2	0.20	0.15		

Standard errors in parentheses (). Incidence ratios [e^{β}] in square brackets. Omits 1989.
* Statistically significant at 5 percent.
† Statistically significant at 10 percent.
[a] This column drops six states (ID, MT, ND, NH, SD, and WY) because no new plants located in those states during any of the 16 years.
[b] This column drops 10 states (those in column (3) plus NE, NM, RI, and UT) because no new plants located in those states during any of the 16 years.
[c] SIC codes 26, 28, 29, 32, 33, 34, 37. (See footnote 15.)

Table 6
Alternative Count Data Models
Of New Foreign-owned Plants
As a Function of Abatement Costs
1977 − 1994

	Stringency measures	Pooled		Fixed-Effects	
		All Manufacturing (1)	Polluting industries (2)	All manufacturing[a] (3)	Polluting industries[b,c] (4)
(1)	Industry-adjusted index (S^*) using negative binomial model.	−0.066 (0.199)	−0.029 (0.237)	0.381 (0.281)	0.636[†] (0.382)
(2)	Unadjusted index (S) of abatement costs using Poisson.	3.53 (4.45)	15.3* (5.6)	46.7* (15.6)	69.1* (20.1)
(3)	Unadjusted index (S) using negative binomial model.	−1.54 (7.10)	11.67 (7.93)	44.8* (16.5)	64.2* (20.2)
(4)	Five-year averages, using industry-adjusted index (S^*).[d]	−0.577* (0.155)	−0.588* (0.213)	−0.211 (0.490)	0.391 (0.681)
(5)	Five-year averages, using unadjusted index (S).[d]	−11.3* (5.3)	0.95 (6.66)	42.7 (26.0)	78.2* (32.7)
(6)	Zero-inflated Poisson, using industry-adjusted index (S^*)	−0.171 (0.159)	−0.053 (0.238)	--	--
(7)	Zero-inflated Poisson, using unadjusted index (S).	3.48 (5.47)	15.53* (7.03)	--	--
	n	768	768	672	608

Standard errors in parentheses. Omits 1989, when no new plant data were collected.
* Statistically significant at 5 percent.
[†] Statistically significant at 10 percent.
[a] The fixed effects models drop ID, MT, ND, NH, SD, and WY because no new plants located in those states during any of the 17 years.
[b] This column drops 10 states (those in column (3) plus NE, NM, RI, and UT) because no new plants located in those states during any of the 16 years.
[c] SIC codes 26, 28, 29, 32, 33, 34, 37. (See footnote 15.)
[d] Rows (4) and (5) average all dependent and independent variables for 1977-1981, 1982-1986, and 1988-1994, and treats each period as one observation. There are thus 48 states and three periods, for 144 total observations. The last period, 1988-1994, takes a six-year average, because there are no new-plant data for 1989.

Data Appendix

Gross Value of Property, Plant and Equipment (PP&E) of Foreign-Owned Manufacturers

Bureau of Economic Analysis (BEA), U.S. Department of Commerce, *Foreign Direct Investment in the United States.*

Employment of Foreign-Owned Affiliates

Bureau of Economic Analysis (BEA), U.S. Department of Commerce, series *Foreign Direct Investment in the United States.*

New Foreign-owned manufacturing plants

International Trade Administration (ITA), Department of Commerce. These data were culled from generally available public sources, transaction participants, and a variety of knowledgeable contacts. The major portion of the data were derived from public secondary sources such as newspapers, magazines, and business and trade journals, as well as from the public files of Federal regulatory agencies.

The data contain the country of origin of the investment, the name of the business enterprise, the 4-digit SIC code of the business enterprise, the reported value of the investment, the state in which the investment was made, the year, and the investment type. Types of foreign direct investment include acquisitions and mergers, joint ventures, real estate transactions, new plants, plant expansions, and equity increases. Any other transaction classified as foreign direct investment is collected under the heading of "other." The Office of Trade and Economic Analysis maintains that the monitoring program identifies the vast majority of significant foreign direct investment transactions in the United States.

New Plant Data: Data on new plants include the state in which the plant was built, the country of origination, the year, the amount of the investment, and the SIC code. We focus on the Manufacturing sector.

Pollution Abatement Costs and Expenditures (PACE) Data

All PACE data were manually entered from tables published by the US Department of Commerce, Bureau of the Census. The variable of interest from this source was the Pollution Abatement Gross Annual Cost (GAC) total across all media types. These data are published in Current Industrial Reports: Pollution Abatement Costs and Expenditures, MA-200, various years. The 1977 data are only for establishments with 20 or more employees. Although survey data was collected from all establishments for the years 1973-1979, in order to lessen the administrative burden on small businesses, they were dropped from the survey, starting in 1980. The PACE Survey was not collected in 1987. Note: There were some censored observations for the state totals.

Gross State Product data:

All gross state product data were acquired via the Regional Economic Information System CD, 1969-1994 published by the US Department of Commerce, Bureau of Economic Analysis, Regional Economic Measurement Division.

Population

Source: Current Population Survey: www.census.gov/population/estimates. Files st9097t1.txt, st8090ts.txt, st7080tx.txt.

Market proximity

This is a measure of how near each state is to potential markets in other states. It is a distance-weighted measure of Gross State Product:

$$M_{it} = \sum_{j \neq i} \frac{Y_{jt}}{d_{ij}}$$

where Y_{jt} is the GSP of state j at time t, and d_ij is the distance from state i to state j (miles between populations-weighted state centroids). Source: BEA. Distances are approximated as a straight line along a great-circle route.

Unionization Rates

Union Membership as Percent of Civilian Labor Force. The Bureau of National Affairs, Inc., "Union Membership and Earnings Data Book: Compilations from the Current Population Survey." Notes: BNA's series begins in 1983. All of the data were obtained through the Statistical Abstracts, except for 1985, 1988, 1990, and 1993, which were obtained directly from BNA. Unionization rates prior to 1983 have been extrapolated from the 1983-1994 trend.

Unemployment

Total Unemployed as Percent of Civilian Labor Force. Source: US Bureau of Labor Statistics, "Geographic Profile of Employment and Unemployment," annual.

Wages

Production Workers in Manufacturing Industries – Average Hourly Earnings by State. Source: US Bureau of Labor Statistics, "Employment and Earnings," monthly. Notes: missing production workers average hourly wages for 1981, and for years prior to 1980. These numbers are interpolated and extrapolated in the data.

Road Mileage

This is the sum of Urban Highway Mileage and Rural Highway Mileage.
Sources: Federal Highway Administration, U.S. Department of Transportation.
www.fhwa.dot.gov/ohim/summary95/section5.html, file hm210.xlw.

Energy Prices

Prices of Fuel and Electricity for Industrial Sector. Source: State Energy Price and Expenditure Report, U.S. Energy Information Administration, www.eia.doe.gov/emeu/sep/, file allprice.csv.

Land Prices

Land Value per Acre. US Department of Agriculture, Economic Research Service, www.econ.ag.gov/Prodsrvs/dp-lwc.htm#prices.

Tax effort

Advisory Commission on Intergovernmental Relations, 1988, State Fiscal Capacity and Effort. This variable measures the extent to which a state utilizes its available tax bases. It is a state's actual revenues divided by its estimated capacity to raise revenues based on a model tax code, multiplied by 100. The national average is 100.