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Hugh Ward[1] and Xun Cao[1]

Abstract

Economists have made a strong case for the relative efficiency of market-based mechanisms for environmental regulation such as cap and trade and "green taxes," yet the spread of these forms has been limited, and traditional "command and control" regulation still predominates. The authors explain geographical and temporal variation in green tax burdens by considering their domestic and international determinants, modeling international influences using spatial lags. Hypotheses are tested using panel data on Organisation for Economic Co-operation and Development member states from 1995 to 2004. At the domestic level, the authors show that green tax burdens are influenced by the left–right and environmental positions of legislative medians and the power of the energy-producing sector, among other factors. The authors also show that ideas about policy diffuse through international networks generated by trade and environmental intergovernmental organizations, but they do not find compelling evidence for international tax competition.

Keywords

green taxation, special interests, legislative medians, policy diffusion, spatial models

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Since Pigou (1920) argued that a tax should be imposed in such a way that polluters' actions reflect their full social cost, economists have come increasingly to recognize the advantages of market-based forms of environmental regulation. Pollution taxes are one of a range of such incentive-based methods, including cap-and-trade emissions trading (Andersen, 1994; Fujiwara, Núñez Ferrer, & Egenhofer, 2006; Organisation for Economic Co-operation and Development [OECD], 1999, 2008a; Stavins, 2003). Since the 1980s the OECD has promoted market-based environmental regulation on efficiency grounds (OECD, 1989), but there is still considerable variation across members in tax burdens. We show that tax burdens are influenced by the positions of legislative medians and the power of the energy-producing sector, among other factors. We show that they are also influenced by the diffusion of ideas through international networks generated by trade and environmental intergovernmental organizations (IGOs), but we do not find compelling evidence for international tax competition.

Many pollution taxes are piggybacked onto existing taxes on energy consumption and motor transport, both for administrative convenience and because this sometimes reduces political opposition (OECD, 2008a). They are often set too low to fully internalize environmental social costs, fail to be set at the uniform level on all equivalent polluting activities (required to achieve efficiency), and are often not paid at all by members of powerful lobbies (Ekins & Speck, 1999; Jordan, Wurzel, & Brückner, 2003; OECD, 2008a; Stavins, 2003). Thus, they can be hard to distinguish from measures aimed purely at raising revenue. Nevertheless, the OECD defines "environmentally related taxes" as

> any compulsory, unrequited payment to general government levied on tax bases deemed to be of particular environmental relevance. Taxes are unrequited in the sense that benefits provided by government to taxpayers are not normally in proportion to their payments. (OECD, 1999, p. 56)

In defining such taxes,

> The name, or the expressed purpose, of a given tax is not a criterion in this project. The focus is instead on the potential environmental effects of the given tax, which is determined by the impacts of the tax on the producer and consumer prices in question, in conjunction with the relevant price elasticities. (OECD, 1999, p. 57)

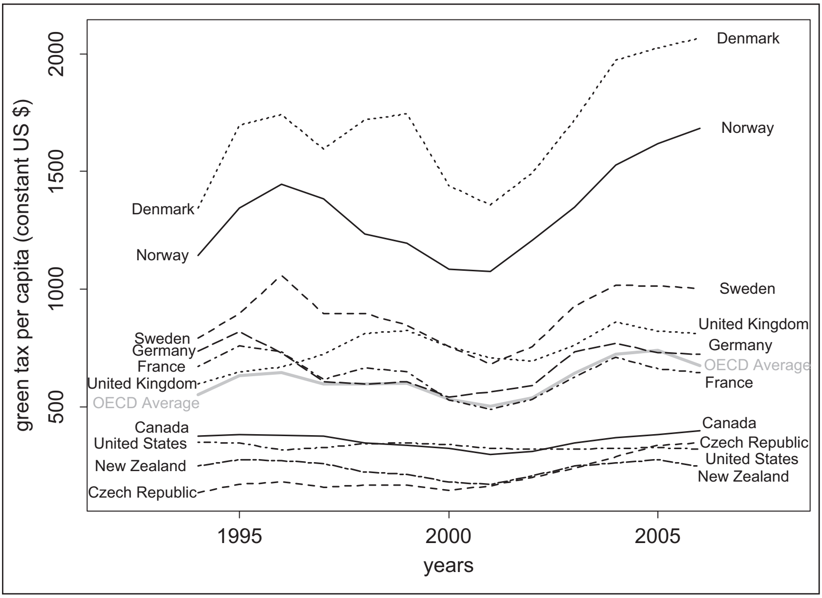

Figure 1. Green taxation burdens in selected OECD countries, 1994–2006

In this article, we use OECD data on what for brevity we call green taxes. We focus on variation across OECD member states over the period 1994 to 2006 in green taxes per capita, denominated in year 2000 U.S. dollars (*green tax per capita*).[1] In 2005, per capita green taxation averaged $741.5 over the 30 OECD member states (Standard Deviation = $532.9), whereas burdens were as high as $2,116.5 in Luxembourg, $2,026.4 in Denmark, and $1,324.3 in the Netherlands. Thus, green taxes generate nontrivial burdens. Figure 1 shows the evolution of per capita green taxation from 1994 to 2006 for selected OECD countries. Although burdens are generally lower in poorer member states like the former communist countries, they are also relatively low in some richer countries like the United States. There is also a degree of coordination through time, especially among the countries with higher tax burdens. These two features of the data suggest that although domestic differences will explain some of the variance, states' tax policies also react to each other.

Traditional environmental regulation has taken such forms as prohibiting certain activities, setting limits on emissions, placing safeguards so that ambient standards in some geographically defined area are maintained, and mandating some technological approach to pollution control. The idea that these "command and control" methods are less efficient than market-based regulation is in common currency among economists,[2] although there is a range of widely recognized practical difficulties (OECD, 2008a; Stavins, 2003). In contrast, the case for market instruments has not been found compelling by parts of the green movement: They dislike rich people and countries buying the right to pollute when green taxes hurt poorer people who cannot easily avoid paying; and (deeper) greens dislike the idea of marketizing nature and basing efficiency criteria purely on human welfare, preferring legislative measures aimed at directly outlawing or limiting certain practices (Beder, 1996; Goodin, 1994). We note, though, that there is some cross-national variation; for example, environmental nongovernmental organizations in the United States appear to have come around (Stavins, 2007). Given political resistance to higher taxes in many countries and the green movement's suspicions, green taxation is one of the most likely areas of environmental politics to observe the impact of voter reaction. For instance, in 1993 the United Kingdom introduced annual fuel excise duty increases of 3% per-annum. But this was withdrawn after protests in 2000 amid accusations that taxes were too small to affect relatively inelastic demand and were merely a revenue-raising device.

Our general theoretical approach suggests that environmental policy is determined by domestic political factors, the policies of other states through policy diffusion mechanisms, and the ways in which domestic politics conditions the impact of international factors.[3] Several studies in the recent literature on comparative environmental political economy also take into account domestic factors and cross-national policy diffusion mechanisms (Bernauer, Kalbhenn, Koubi, & Ruoff, 2010; Holzinger, Knill, & Sommerer, 2008; Jahn, 2009; Knill, Debus, & Heichel, 2010; Perkins & Neumayer, 2008). We advance this line of research by taking the first step to model how domestic politics condition policy diffusion: A novel feature of our approach is that we test the hypothesis that the ideological affinities between key decision makers condition policy learning between states. More importantly, our study is the first in the political science literature to focus on green taxation using large-*N* methods.[4] We therefore believe it is better to adopt a broad approach, considering a variety of possible influences, rather than focus narrowly on one factor and risk biased inference resulting from underspecified models.

Much effort has been put into conceptualizing causal mechanisms behind international policy diffusion, including competition, learning, emulation, and coercion (Dolowitz & Marsh, 2000; Franzese & Hays, 2008; Simmons, Dobbin, & Garrett, 2006).[5] Of particular interest to us are studies applying these ideas to environmental policy diffusion (Holzinger et al., 2008; Tews, Busch, & Jörgens, 2003). Recent research on environmental policy diffusion sometimes uses spatial lags to represent international influences (Jahn, 2009; Perkins & Neumayer, 2008). To help disentangle correlated international diffusion processes we use multiparametric spatiotemporal autoregressive (m-STAR) models, which are now coming into use in the literature (Hays, Kachi, & Franzese, 2010).

The Political Economy of Market-Based Environmental Regulation

In this section we discuss some key domestic and international influences on environmental policy. Some additional factors are considered when we discuss control variables in the section on measurement.

Domestic Influences on Tax Burdens

Special interests. Oates and Portney (2003) argue that the best way to understand environmental regulation is through a model in which interest groups contend to influence policymakers. In a simple lobbying model where the government maximizes a convex combination of the utility functions of an industry and an environmental group over environmental quality standard and how market-orientated the form of regulation is, typically the standard will be too lax, and too little use will be made of market-based regulation because of the preferences of industry for command and control (Fredriksson, 1997; Hahn, 1990). Within a "menu-auction" framework for understanding lobbying,[6] if all groups organize fully, policy will be socially efficient; so inefficient policy is the result of differential failure to organize (Aidt, 1998). This is likely to occur because there are generally concentrated costs for producers and dispersed benefits for citizens, leading to inefficient exemptions and subsidies (Felder & Schleiniger, 2002).[7] The continuing predominance of emission standards over charges may be explained in these terms when it is recognized that standards generate barriers to entry enjoyed by established firms (Buchanan & Tullock, 1975), whereas regulatory bureaucracy's and Congress' general preference for command and control (Stavins, 2007) might be explained in terms of agency capture and political donations. In contrast, Aidt and Dutta's (2004) model suggests that as environmental standards get tighter, both citizens and polluting industries move away from support for inefficient command and control, but a conflict emerges between them over whether green taxes or tradable permits will be used because voters get benefits from recycled revenues from green taxes whereas polluters benefit from "grandfathered" permits. The way the conflict is resolved depends on politicians optimizing support.

In summary, the literature leads us to expect that green tax burdens will fall with the power of energy producers and energy intensive sectors: Generally, they will try to block measures that limit their use of energy; specifically, if market-based measures are contemplated, they will push for grandfathered permits rather than green taxation.

Legislative medians. Legislative median positions are likely to be an important constraint on environmental policy change (Jahn, 2009). Allowing for the weight given to constituents' interests by legislators, legislative voting over the degree of environmental regulation and the type of regulation used may be inefficient because of the induced preferences of the median legislator (Campos, 1989). If there is an association between stakeholders in an industry and parties, and if the party of a high-polluting industry is in power, taxes need not be set efficiently high; and because of fluctuations deriving from which party is in power, the lower variance of command and control may even be more efficient (Boyer & Laffont, 1999).

The idea that median positions matter is consistent with the literature that focuses on party positions, as legislators' positions can be expected to correlate by party. Based on a study of Scandinavia, Daugsberg and Svendsen (2001) argue that social democratic parties are more likely to adopt green taxes because of the reluctance of parties of the center-right to tax producer interests. Center parties may well oppose green taxes, too, because of links with specific groups like farmers. The evidence from large-*N* studies is mixed: Although some authors find that the dominance of the left correlates with improved performance on pollution indicators (Jahn, 1998; Neumayer, 2003), others find that this variable does not correlate significantly with environmental policy output but that party and government positions on the environmental dimension matter (Knill et al., 2010; Spoon & Jensen, 2009).

Although the politics of green taxation is potentially multidimensional, raising the specter of chaotically unstable legislative majorities, we argue that there is institutional separation between the dimensions, generating institutionally induced equilibria (Shepsle, 1979): Whether to regulate polluters using green taxes is a question for environmental ministries, ministers, specialist committees, and bills, whereas decisions about the tax rate, exemptions, and how revenue is recycled are strongly influenced by ministries of finance and associated legislative organs because green taxes are most frequently imposed through higher rates on existing taxes.

Considering one dimension, with institutional issue separation the spatial model suggests that change of the status quo policy on that dimension is governed by the winset of the status quo. In turn, this is governed by the position of the median member of the legislature relative to the status quo. Consider the tax rate dimension. Suppose the status quo tax rate is lower than the median's ideal. The median and those who prefer higher tax rates form a majority; but the greatest tax increase that can gain the support of this majority is governed by the member most loath to see an increase—the median. If the political right is less prone to support green taxation, *conditional on the status quo tax rate*, the further to the right the legislative median, the lower the equilibrium tax rate, if we assume an open agenda.[8] On the environmental dimension, the simplest hypothesis is that the more proenvironment the legislative median, *conditional on the status quo tax policy*, the higher the green tax rate. However, as we have seen, some environmentalists and deep green parties are against market-based environmental regulation, including green taxes (Beder, 1996; Goodin, 1994). We therefore suspect that there may be nonlinearities between the environmental emphasis of the legislative median and the level of green taxes: The more proenvironment the median legislator, the higher the green tax, but this relationship holds up only to moderate emphasis on green themes.

Case studies in Western Europe suggest the importance of coalition–government dynamics for environmental policy (Jordan et al., 2003). Here the situation is complex, considering that formation of coalitions, allocation of cabinet portfolios, and decisions over policy might be bargained interdependently. Competing models exist, and our understanding of the consequences for equilibrium spatial decision making over policy is currently limited (Austen-Smith & Banks, 2005; Diermeir, 2006; Laver & Shepsle, 1996). We assume that constraints on changing the status quo are governed by the position of the median member of the cabinet relative to the status quo. Below we note that our operationalization of this point is consistent with some empirical evidence.

International Influences on Tax Burdens

International tax competition. Daly (2001) predicts a race to the bottom in environmental regulation. However, the evidence for this from case studies is mixed (Clapp, 2001; Holzinger & Knill, 2004; Wheeler, 2001), suggesting that such races are conditioned by domestic demand, institutional path dependence, and administrative capacity (Jordan et al., 2003; Kern, Jörgens, & Jänicke, 2001). Where a country with a large domestic market legislates tighter environmental product standards, possibly as a result of powerful domestic producers' protectionist lobbying, international competition may drive standards up, as foreign competitors vie for access (Vogel, 1995). Freer trade may motivate environmental groups successfully to lobby for tighter regulation, leading them to ally with producers seeking protection (Vogel, 1995).[9] Green taxes may well not be an important factor in location decisions of corporations that trade internationally, as many of them are aimed at consumers, may be passed on to consumers, or are too low to offset more compelling influences on location decisions.

The proposition that green tax rates will be subject to downward pressure because of competition has not been tested in a large-*N* study, so we focus on parallel debates on other taxes. Here empirical work suggests that first-generation theory about races to the bottom (Wilson, 1986) ignored the way competition is conditional on domestic political institutions (e.g., partisan politics and electoral system), government budgetary constraints, and fairness norms (Plumper, Troeger, & Winner, 2009; Swank & Steinmo, 2002). We control for a number of domestic factors that might condition competition over green taxes. We also expect the level of tax competition between any two countries to be proportional to the level of competitive pressure between them, which is ultimately a positive function of how similar or substitutable they are in the networks of trade (Burt, 1992).

Information, policy learning, and green taxation. For policy diffusion, what might appear to be competitive races in taxation might better be seen as the consequence of diffusion of neoliberal ideas about good practice in reducing tax burdens. Although diffusion may occur through imposition by powerful states or international institutions, as regards environmental policy, Busch and Jörgens (2005) argue that harmonization through international regime bargaining is more plausible than imposition. Such harmonization may offset competitive races to the bottom, in the case of the EU (Holzinger & Knill, 2004). Front-runner states, as well as international organizations and transnational networks, may be important to harmonization upward in environmental regulation through learning (Kern et al., 2001; Tews et al., 2003). Case studies suggest that front-runners on carbon taxes have generally been richer Western European and, especially, Nordic states (Tews et al., 2003, p. 587).

A possible diffusion mechanism is that states with less technical and administrative capacity learn from the existing policies of states with more. We allow for the potential influence of more developed OECD states, some of which are front-runners on green tax policy, on poorer members. We also propose that countries with close network ties through trade and membership of international organizations are likely to influence each other's policies on green taxation. In general, the closer the two countries are in international networks, the more similar they are in terms of domestic economic policies because close connections facilitate policy learning and therefore policy diffusion.

Policy Learning Conditioned by Ideological Affinity

Our theoretical framework focuses on domestic political factors, the policies of other states through policy diffusion mechanisms, and the ways in which domestic politics conditions the impact of international factors. In this article, as a first step, we theorize and test only one particular conditional effect: policy learning conditioned by ideological affinity of legislative medians. Little theoretical consideration has been given in the literature on diffusion to the conditions under which information transfer will occur, yet there is an abundant literature on this in relation to domestic lobbying, based on games of asymmetric, incomplete information where *j* knows something that *i* does not about states of the world that condition which policy is best. The key result is that, in the most informative equilibria, credible transmission of information depends on how close *j* and *i*'s preferences are, that is, successful lobbying is likely to be between those with similar preferences because of distrust of those with dissimilar preferences (Austen-Smith, 1997; Potters & van Winden, 1992). Countries may be close ideologically without being close geographically. If key legislative players in both countries have similar preferences, more information is likely to be transferred between them. Connecting with our previous discussion, we consider two dimensions of underlying similarities in government preferences, one on environmental emphasis and the other on left–right partisan center of gravity. We expect a higher level of information transfer and policy diffusion between governments whose legislative medians share similar ideological preferences.

Methodological and Measurement Issues in Modeling Domestic and International Influences

Modeling international influences requires us to use estimation techniques that explicitly incorporate mechanisms of policy interdependence. The spatial lag model is often considered conceptually appropriate because the dependent variable for a unit is better conceptualized as being affected by the values of dependent variable in nearby units. In a spatial lag model that also includes a temporal lag, or a STAR model (spatial temporal autoregressive model),

$$\mathbf{y_t} = \varphi\, \mathbf{y_{t-1}} + \mathbf{X_t}\, \beta + \rho \mathbf{W} \mathbf{y_t}, \tag{1}$$

Y is an $NT \times 1$ vector of observations, for N countries over T time periods,[10] and **X** the battery of domestic variables. We include country fixed effects in **X** to capture idiosyncratic national path dependencies in taxation policy and other forms of cross-sectional heterogeneity. The connectivity matrix, **W**, is $NT \times NT$ matrix with $TN \times N$ submatrices along the block diagonal, and typical element $w_{i,j,t}$ capturing relative connectivity or influence from unit j to i at time t. $\mathbf{W}\mathbf{y_t}$ is the spatial lag: It gives a weighted average of other observations in the year concerned, with each weight specified by $w_{i,j,t}$. In Equation 1, the estimation of ρ (spatial coefficient) captures the strength of policy interdependence. We add the time-lagged dependent variable $\mathbf{y_{t-1}}$ to capture dependency in time, and also because our hypotheses about the effects of legislative vetoes are contingent on the status quo, which we take as last year's tax policy.[11]

The problem with estimating Equation 1 using ordinary least squares (OLS) is that the dependent variable is on both sides of the equation, which causes simultaneity bias (Franzese & Hays, 2008). One simple strategy to get around is to time lag the spatial lag by 1 year and use OLS estimation. This is only a solution to simultaneity bias if the errors, $\varepsilon_{i,t}$, are serially independent (Franzese & Hays, 2007a, 2008). Including a temporally lagged dependent variable often helps to eliminate serial correlation in the errors, but there is no guarantee. More important, the assumption in this approach is that it takes a time lag to affect the policy outcome in another connected country. With budgetary cycles coordinated across most OECD countries, to assume that there can be no effect within the year concerned seems to be unjustified. Moreover, even in the best situation where this assumption holds, as Franzese and Hays (2007b) have pointed out, the OLS model with time-lagged spatial lags provides unbiased estimates only if the first observations are nonstochastic, that is, they are fixed across repeated samples. In our case, the spatial maximum likelihood approach is suggested as it does not assume a temporally lagged spatial lag and addresses simultaneity bias head on (Elhorst, 2001; Franzese & Hays, 2008).

Measuring Domestic Political Determinants of Green Taxation

Domestic lobbies' demand. We expect tax burdens to fall with the power of energy intensive sectors and energy producers. Energy intensive sectors can be identified, for instance by reference to those included in the EU's CO_2 emissions trading scheme such as electricity generation, cement production, and glass making. In principle, the power of such sectors could be measured by their contribution to GDP, but available breakdowns of GDP are not fine enough to make this practicable. Things are more straightforward on the production side. We used World Bank World Development Indicators (WDI) data on national energy production in kilotons of oil equivalent and divided by real GDP.[12] This gives energy production per unit of real gross domestic product (*energy production per GDP*). Essentially similar results to those reported in the text were obtained from WDI data on kilograms of CO_2 emitted per dollars of GDP at 2005 purchasing power parity.[13]

Legislative medians. Because cross-nationally comparable individual-level data do not exist on individual legislators' positions, we need to infer the position of the median from data on party positions. Given a party's positions on a relevant dimension, we assume that its legislators are uniformly distributed over an interval centered on the party position and delimited by the cut points halfway between that position and those of adjacent parties.[14] If no party or coalition held an overall majority, we operationalize the position of the median as the average of all parties' positions, weighting by number of seats.[15] If there was a single-party majority government, we use that party's position. If there was a majority coalition, we use the weighted average by seat shares in the coalition of members of the coalition, which is consistent with the evidence that this is the most important determinant of coalition policy (Warwick, 2001). *Legislative median: Left–right* is the left– right dimension median calculated in this way; and *legislative median: environmental* is the median's position on the environmental dimension.[16]

In carrying out these calculations we used data from the Comparative Manifestoes Project (CMP) on parties' right–left scores and the emphasis they placed on the environment (Klingemann, Volkens, Bara, & Budge, 2006). The CMP content analyses quasi-sentences. Left–right scores are based on the percentage of sentences that are held to reflect themes of the ideological right minus the percentage held to reflect themes of the ideological left. Environmental scores are based on the percentage of sentences mentioning the environment, as on this valence issue negative comment was largely absent during the period the data cover. For the period up to the last national election before 2000, we used Cusack and Fuch's (n.d.) data set, tying the CMP data to data on governing coalitions. For the remaining years up to 2004, when Cusack and Fuch's data set ceases to be available, we extracted information about which coalitions formed from the Inter-Parliamentary Union's Parline database. We assume parties' left–right scores do not change between elections. In election years or years when there was a breakdown of the coalition, we weight scores proportionally to the number of months before the election or coalition shift.

International Influences Captured by Spatial Weights

In Equation 1, the connectivity matrix **W** captures relative connectivity or influence from unit j to i at time t. The spatial lag $\mathbf{Wy_t}$ is the weighted average of other observations in a given year, with each weight specified by $w_{i,j,t}$. We row standardize each connectivity matrix so that the estimated values of ρ reflect the average influence of other countries' tax burdens. For presentational simplicity we discuss definition of weights prior to row standardization. Some weights are time dependent, but we drop time subscripts when defining lags.

Competition for export market. Competition induces policy interdependence among countries that aim at the same export markets. A country's export profile is composed of $k\times(n\text{-}1)$ elements where $(n-1)$ is the number of potential export destination countries and k is the number of trade sectors. The intensity of competition between countries can be captured by their level of structural similarity—the correlation between their export profiles (Smith & White, 1992). Using the UN's Standard International Trade Classification (SITC), we identified 10 broad sectors, defined by one-digit SITC.[17] Structural equivalence is bounded between –1 and 1, but it is only when countries sell similar products in similar markets that they compete. So we define the connectivity matrix $\mathbf{W}^{structural\ equivalence}$ such that the elements, $w_{i,j}^{structural\ equivalence} = w_{j,i}^{structural\ equivalence}$, is the correlation between export profiles of country j and i if this correlation is positive, else set to 0.[18]

Coaffiliation to environmental IGOs. We anticipate higher levels of policy interdependence between countries that are more connected to each other through coaffiliation to environmental IGOs.[19] So in the connectivity matrix-$\mathbf{W}^{environmental.IGOs}$, each element ($w_{i,j}^{environmental.IGOs}$) is the number of common environmental IGO memberships shared by country i and j. We used Ingram's data on environmental IGOs and chose to use his minimalist category because these IGOs have no coercive power over member states, excluding coercion as an alternative mechanism.[20] We expect a positive spatial coefficient ρ.

Dyadic trade. A country is likely to learn more from its major trading partners; therefore, we define connectivity matrix $\mathbf{W}^{dyadic.trade}$ in which $w_{i,j}^{dyadic.trade}$ is the total trade flows from i to j plus total trade flows from j to i, as reported in the Correlates of War trade data Version 2.0 (Barbieri, Keshk, & Pollins, 2008).

Diffusion from more developed economies. We argued earlier that ideas about green taxes may diffuse from richer to poorer countries. So we define $\mathbf{W}^{diff:GDP.per.cap}$ in which $w_{i,j}^{diff:GDP.per.cap} = Real\ GDP\ per\ capita_j - Real$

GDP per capita$_i$ if *Real GDP per capita*$_j$ > *Real GDP per capita*$_i$, and 0 otherwise.

Geographical distance. Finally, we control for the effects of geography using the connectivity matrix $\mathbf{W}^{distance}$ such that $w_{i,j}^{distance} = w_{j,i}^{distance}$ = the distance between capital cities of country *i* and *j* in kilometers, as calculated by Gleditsch.[21] We expect a negative ρ.

Conditional Effect of Ideological Affinity on Learning Captured by Spatial Weights

We argued above that flows of information will be conditioned by the degree of ideological similarity across pairs of countries. We use spatial weights to capture the conditional effect of ideological affinity for policy learning.[22] First, we defined $\mathbf{W}^{diff:left\text{–}right}$ such that each element $w_{i,j}^{diff:left\text{–}right}$ represents the ideological difference between country *i* and *j*, which is measured as the absolute difference between the positions of the legislative medians on the right–left dimension: $w_{i,j}^{diff:left\text{–}right} = w_{j,i}^{diff:left\text{–}right} = |\, LegislativeVeto{:}left.right_i - LegislativeVeto{:}left.right_j|$. Second, we define $\mathbf{W}^{diff:environmental}$ in similar fashion: $w_{i,j}^{diff:environmental} = w_{j,i}^{diff:environmental} = |LegislativeVeto{:}environmental_i - LegislativeVeto{:}environmental_j|$. We expect these weights to have negative ρ values: Information transfer is less likely the greater the ideological distance.

Other Socioeconomic Variables Included as Controls

It is common to include green party vote share or membership of the government in the literature (Holzinger et al., 2008; Spoon & Jensen, 2009). Our estimated models include *green party*, which equals 1 if the CMP reports that a green party elected members to the lower house. It is possible that other tax rates influence green taxation, but we do not have strong theoretical expectations. On one hand, the higher other taxes are the greater the resistance may be to green taxation among citizens. On the other hand, citizens may generally be willing to pay higher taxes because they are satisfied with state provision of welfare and other benefits. We include taxes on personal income as a percentage of GDP (*income tax*; OECD, 2008b).[23]

One argument for the environmental Kuznets curve posits that the environment is a relatively low priority for citizens in the early stages of development but it becomes a higher priority as they become better off (Grossman & Krueger, 1995; Seleden & Song, 1994). To control for this nonlinear income effect, we included real GDP per capita (*real GDP per capita*) and its square term (*real GDP per capita squared*). We also control for unemployment (*unemployment rate*) using WDI data. Green taxes, as other types of taxation, generate revenues for governments. When unemployment rates are high, governments are often under pressure to increase taxation to provide unemployment support; therefore, we expect green taxes to increase with the level of unemployment rate. However, fiscal expansion cannot go on forever: When the economy is in deep crisis and business is unwilling to hire, governments might choose to lower the tax burden. In sum, we expect an inverted U-shaped relationship between unemployment rate and per capita green tax, and we test this by including *unemployment rate squared*. Finally, because the business cycle is somewhat coordinated across the OECD, controlling for unemployment also helps to allow for common economic shocks.

Globalization theory suggests not only that a country will respond to competitiveness concerns in relation to trading partners or competitors but also that there may be a more general effect, whereby countries more heavily engaged in the global economy have lower tax rates. Rather than increase complexity of presentation by including measures of separate dimensions of openness, we use Dreher's (2007) index of actual economic flows (*actual economic flows*), derived from principal components analysis of data on a nation's trade flows, foreign direct investment, and portfolio investment.

Empirical Findings

One Spatial Lag Models

Table 1 presents estimates of models with one spatial lag for the OECD countries from 1995 to 2004. For brevity we do not report coefficients on country fixed effects.[24] Results in relation to domestic variables are quite similar across Models 1–7 with different spatial lags and can be discussed together. We focus on variables' short-term impacts, that is, the reported coefficients, not equilibrium effects. As anticipated, *legislative median: left–right* has a significant negative coefficient.[25] This confirms our expectation that the further to the right the legislative medians, the lower the level of green taxes. We thought the impact of the legislative median position on the environmental dimension (*legislative median: environmental*) might not be linear. Across models in Table 1 there is a definite quadratic relationship with maxima around 4—just below the mean score in the sample of around 4.6. The distribution of *legislative median: environmental* is quite skewed, with relatively few cases above the mean value. So one interpretation of the nonlinear relationship is that the short-run impact on green tax burden is highest with a

Table 1. Models With One Spatial Lag

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
	$\mathbf{W}^{distance}$	$\mathbf{W}^{environmental.IGOs}$	$\mathbf{W}^{dyadic.trade}$	$\mathbf{W}^{structural.equivalence}$	$\mathbf{W}^{diff:GDP.per.cap}$	$\mathbf{W}^{diff:left\text{–}right}$	$\mathbf{W}^{diff:environmental}$
Green tax per capita lagged	0.390 (9.95)**	0.383 (10.18)**	0.386 (10.58)**	0.409 (10.51)**	0.394 (10.08)**	0.390 (10.22)**	0.392 (10.09)**
Legislative median: left–right	–1.100 (2.38)*	–0.969 (2.18)*	–0.870 (2.01)*	–1.159 (2.51)*	–1.063 (2.30)*	–1.298 (2.86)**	–1.100 (2.40)*
Legislative median: environmental	18.499 (2.68)**	16.989 (2.56)*	15.764 (2.44)*	15.306 (2.20)*	17.516 (2.54)*	16.760 (2.48)*	30.628 (4.01)**
Legislative median: environmental, squared	–2.223 (3.29)**	–2.049 (3.16)**	–2.061 (3.26)**	–1.894 (2.79)**	–2.113 (3.13)**	–1.998 (3.03)**	–3.036 (4.26)**
Green party	20.067 (1.76)	23.045 (2.10)*	19.865 (1.86)	21.114 (1.85)	18.363 (1.61)	19.409 (1.74)	21.704 (1.91)
Real GDP per capita	0.001 (0.09)	–0.005 (0.69)	0.001 (0.20)	0.003 (0.40)	–0.005 (0.64)	0.001 (0.19)	–0.002 (0.32)
Real GDP per capita squared	0.000 (1.13)	0.000 (1.69)	0.000 (1.38)	0.000 (0.22)	0.000 (1.57)	0.000 (1.16)	0.000 (1.51)
Unemployment rate	27.904 (4.29)**	22.141 (3.48)**	22.964 (3.77)**	29.505 (4.61)**	27.681 (4.25)**	25.183 (3.95)**	25.650 (3.91)**
Unemployment rate squared	–1.011 (4.37)**	–0.843 (3.75)**	–0.835 (3.85)**	–1.070 (4.70)**	–1.006 (4.35)**	–0.913 (4.02)**	–0.932 (4.00)**
Actual economic flows	–2.452 (3.47)**	–1.978 (2.93)**	–1.640 (2.48)*	–2.903 (4.33)**	–2.358 (3.31)**	–2.075 (3.00)**	–2.341 (3.35)**
Energy production per GDP	–284,661 (3.45)**	–219,047 (2.73)**	–248,012 (3.26)**	–279,021 (3.38)**	–273,454 (3.29)**	–242,718 (2.99)**	–276,187 (3.38)**
Income tax	1.192 (12.58)**	1.088 (11.77)**	0.998 (10.70)**	1.251 (14.05)**	1.156 (11.73)**	1.122 (11.99)**	1.169 (12.42)**
Constant	27.828 (0.21)	28.844 (0.23)	–176.627 (1.39)	58.331 (0.45)	147.962 (1.16)	–55.032 (0.42)	47.789 (0.37)
Spatial Coefficient: ρ	.219 (3.02)**	.310 (5.65)**	.432 (6.83)**	.112 (3.28)**	.098 (3.16)**	.284 (4.73)**	.186 (3.65)**
Observations	266	266	266	266	266	266	266
Moran's I	–.058 (2.387)*	.033 (2.148)*	.105 (3.701)**	.015 (0.997)	.189 (6.335)**	.005 (0.471)	–.055 (2.664)**
Robust LM stat.	9.007**	20.914**	22.426**	4.548*	6.287*	11.898**	8.936**

Fixed effects not reported. Absolute values of *t* or *z* statistics in parentheses.
$*p < .05$. $**p < .01$.

moderate emphasis on the environment but falls in the relatively infrequent cases where the legislature sees the issue from a deeper green perspective. We find that *energy production per GDP* has a significant negative coefficient across models. Although we have expressed caution about whether this fully captures the impact of powerful energy lobbies, it tends to support the view that they constrain the level of green taxation.

Green party has a positive coefficient but achieves significance only in Model 2. The estimated coefficients for *real GDP per capita* and its square term in Table 1, however, provide no significant evidence for environmental Kuznets curve between income and green tax. On the other hand, we find a strongly nonlinear relationship between *unemployment rate* and the green tax: Models in Table 1 report a significant inverted U-shaped relationship. The interpretation of this finding also supports our expectation that highlights the role of green taxes as means of public finance: Governments tend to increase taxation to finance public spending for increased unemployment until the point when the economy is in deep crisis. In the sample, the range of unemployment is from 1.8% to 23.9%, with a mean around 7.5%. With a maximum estimated at around 14% unemployment for the inverted U-shaped relationship, this suggests that when unemployment is substantially above the mean, governments tend to push down green tax burden for the economy.

As expected, *actual economic flows* generally has a significant negative coefficient, indicating that being tightly tied to the world economy has a negative short-run impact on levels of green tax. We find that higher income taxes are associated with higher green taxes. This might indicate that in systems where citizens are generally willing to pay higher taxes they are also willing to pay higher green taxes; further research is needed to establish what lies behind this result.

For each spatial lag used, we report Moran's I, where a significant value suggests clustering (positive values) or dispersion (negative values) of the dependent variable on the lag concerned. We also report a robust Lagrange multiplier (LM) test for the significance of the inclusion of a spatial lag in the OLS estimate of the model (Anselin, Bera, Florax, & Yoon, 1996). Across all the models the LM test indicates that including a spatial lag is justifiable. Model 1 includes a spatial lag based on distance between capital cities ($\mathbf{W}^{distance}$). The ML estimate of the spatial coefficient ρ is positive and significant, which is unexpected on theoretical grounds and in the light of the negative value of Moran's I. In Model 2, the spatial lag on shared memberships in environmental IGOs ($\mathbf{W}^{environmental.IGOs}$) has a positive and significant value of ρ, suggesting that countries strongly linked via the environmental IGOs influence each other's tax policies. Both bilateral trade connections ($\mathbf{W}^{dyadic.trade}$)

and structural similarity in exports ($\mathbf{W}^{structural.equivalence}$) have significant positive values for ρ, as expected (Models 3 and 4), although Moran's I for structural similarity is not significant. Model 5 suggests that richer countries positively and significantly influence the policies of poorer countries (positive significant ρ based on $\mathbf{W}^{diff:GDP.per.cap}$). In relation to spatial lags based on similarity of dyads' effective median positions on the right–left ($\mathbf{W}^{diff:left\text{-}right}$) and environmental ($\mathbf{W}^{diff:environmental}$) dimensions, our theoretical expectations were confounded, with positive ρ values on spatial lags reported in Models 6 and 7.

M-STAR Models With Multiple Spatial Lags

Spatial lags in the last section are often quite highly related (see Table 1 of the online appendix), so models in Table 1 are liable to give biased estimates of spatial effects, as the single lag included partly acts as a proxy for others. We addressed this by estimating a series of m-STAR models with different combinations of the seven spatial lags, while keeping the same set of non-spatial-lag variables as in Table 1. Because of space limits, we report these models in the online appendix, where it is seen that allowing for the relationships between spatial lags, the theoretically plausible and statistically significant spatial lags are those based on distance among capital cities ($\mathbf{W}^{distance}$), shared memberships in environmental IGOs ($\mathbf{W}^{environmental.IGOs}$), and bilateral trade ($\mathbf{W}^{dyadic.trade}$). We do not find justification for inclusion of the structural similarity lag ($\mathbf{W}^{structural.equivalence}\mathbf{y}$), ideological distance ($\mathbf{W}^{diff:left\text{-}right}\mathbf{y}$ and $\mathbf{W}^{diff:environmental}\mathbf{y}$), or the lag based on difference in wealth between countries ($\mathbf{W}^{diff:GDP.per.cap}\mathbf{y}$).Model 8 in Table 2 reports the full results for this model, which we feel is the one best supported by the data. In this model, the spatial lag on distance is negative, as we would expect intuitively. Connections in trade and environmental IGOs also induce interdependence on green taxes between countries as suggested by the positive and statistically significant spatial coefficients. Regarding domestic variables, the only notable difference with results in Table 1 is that the *legislative median: left–right* variable is now significant at only the 90% level.[26] In the online appendix, we find that results are robust to alternative ways of conceptualizing and measuring veto structures, interest intermediation, and citizen demand.

Equilibrium Effects

Revenue from taxes on petrol, diesel, and the sale and use of motor vehicles dwarfs that from other green taxes (Fujiwara et al., 2006; OECD, 1999). With only some 14% of greenhouse emissions coming from industrial sources and

Table 2. Final Model

	(8)
Estimation method	m-STAR/ML
Green tax per capita lagged	0.392 (10.96)**
Legislative median: left–right	–0.724 (1.70)
Legislative median: environmental	14.163 (2.24)*
Legislative median: environmental, squared	–1.795 (2.90)**
Green party	22.501 (2.14)*
Real GDP per capita	–0.005 (0.70)
Real GDP per capita squared	0.000 (1.87)
Unemployment rate	20.797 (3.43)**
Unemployment rate squared	–0.793 (3.71)**
Actual economic flows	–1.941 (2.96)**
Energy production per GDP	–236,629 (3.10)**
Income tax	1.012 (11.06)**
ρ: $\mathbf{W}^{distance}Y$	–.481 (3.59)**
ρ: $\mathbf{W}^{environmental.IGOs}Y$	.406 (2.77)**
ρ: $\mathbf{W}^{dyadic.trade}Y$	.346 (2.95)**
Constant	–25.775 (0.19)
Observations	266

Fixed effects not reported. Absolute values of *z* statistics in parentheses.
**p* < .05. ***p* < .01.

24% from power generation (Stern, 2006, p. iv), green taxation is likely to play a vital role in future responses to climate change, with increased taxes on personal transport, home heating and air conditioning, and sectors like agriculture.[27] This makes the estimates derived from our spatial models of the equilibrium effects of tax changes in one country or group on others of particular policy relevance, as they are suggestive of the need for coordinated action if free riding is to be avoided.

With models including spatial lags, coefficients indicate only the short-run impact of a shock to a variable. Nevertheless, it is possible to calculate long-term equilibrium impacts. We carried out the experiment of hypothetically increasing green tax rates in some countries and looking at the long-term effects on all countries, as the shock reverberates through the system of spatial and temporal lags.[28] Based on Model 8, Table 3 reports three such experiments for the impact of a $100 increase in *green tax per capita* in 2003, first for the United States, second for Germany, and third for all EU members simultaneously. The table reports lower and upper bounds of a simulated 90% confidence interval and the median (50%) equilibrium impact, based on

Table 3. Simulated Equilibrium Effects of $100 Increases in Green Tax Burdens in 2003

	$100 increase in U.S. green tax burden			$100 increase in German green tax burden			$100 increase in EU green tax burden		
	5%	50%	95%	5%	50%	95%	5%	50%	95%
United States	120.54	171.50	209.07	3.18	21.48	85.90	−15.79	129.70	538.91
Canada	5.37	59.42	134.11	−8.05	12.83	58.26	−84.02	96.03	445.91
United Kingdom	−555.34	−50.72	6.93	17.64	64.03	390.63	322.60	646.15	2970.44
Ireland	−517.91	−42.76	14.16	17.04	60.16	358.43	312.48	622.01	2758.77
Netherlands	−642.49	−63.31	−1.81	20.03	74.22	431.42	338.43	689.52	3288.89
Belgium	−656.35	−66.12	−3.88	17.74	70.47	431.67	342.21	699.44	3337.04
Luxembourg	−672.48	−68.42	−6.75	21.08	75.80	445.77	341.44	704.40	3341.07
France	−609.09	−56.13	0.78	20.91	71.08	412.73	327.29	666.82	3121.56
Switzerland	−631.74	−62.69	−0.91	17.72	76.53	429.78	184.25	523.16	3069.10
Spain	−587.96	−56.19	−2.81	14.79	63.56	389.53	321.79	649.30	2963.86
Portugal	−563.28	−53.80	−3.13	12.85	59.98	371.16	315.27	631.95	2818.71
Germany	−624.49	−60.47	0.40	165.54	220.89	573.17	332.86	675.81	3212.66
Poland	−627.27	−64.61	−6.39	15.86	78.62	426.46	329.10	666.79	3116.50
Austria	−610.05	−59.99	−3.45	15.44	82.11	425.30	325.12	656.78	3033.68
Hungary	−615.98	−62.44	−4.04	15.95	78.35	421.93	330.26	665.67	3097.41
Czech Republic	−631.77	−63.81	−5.31	14.70	85.07	439.07	331.30	672.85	3167.46
Italy	−587.45	−57.17	−0.24	13.79	65.71	392.58	321.62	647.95	2991.03
Greece	−515.88	−48.16	1.10	11.41	57.54	339.81	302.07	597.76	2578.90
Finland	−545.42	−50.24	1.16	14.94	61.03	364.93	307.71	615.30	2754.50
Sweden	−563.12	−51.26	2.41	16.33	62.63	378.44	311.79	626.21	2884.29
Norway	−573.48	−52.39	1.90	16.46	64.72	387.04	166.69	476.04	2785.48
Denmark	−616.99	−59.95	−2.24	16.17	69.58	412.77	323.79	662.28	3128.31
Iceland	−439.46	−28.60	10.04	10.69	50.51	300.95	124.08	373.69	2118.96
Japan	1.14	35.50	80.53	−2.53	18.83	67.38	−86.89	101.27	427.27
Australia	−17.13	20.03	62.50	−7.07	16.02	65.14	−108.69	94.70	440.81
New Zealand	−14.50	18.64	73.71	−19.93	12.75	58.80	−209.96	75.95	398.87

1,000 random draws from the multivariate normal distribution of the spatial and temporal lags. If the central estimate is significantly different from zero at 90% level, the effect is highlighted. The simulations suggest that a policy shock in the United States would significantly and positively affect Canada and Japan (albeit to a smaller extent than the initial tax change in the United States) but would have significant *negative* effects in many European countries. In contrast, the simulations suggest that Germany could act as an effective leader, a $100 shock there eventually pushing up tax burdens significantly in most European countries and even to a small extent in the United States. Finally, we simulate the effect of a simultaneous $100 shock in all EU countries. Here the effects are quite large across Europe, tax increases in European countries positively feeding back off each other to get much bigger equilibrium increases than the initial shock. We think that these results should give some considerable pause for thought about international action on climate change.

Conclusion and Future Research

Our general theoretical framework emphasizes domestic factors, the influences of other states' policies through policy interdependence, and the interaction between the two. Although we found evidence for the first two factors, we did not find compelling evidence that the ideological affinity between states conditioned the reception of other states' policies. But ideological affinity is only one factor that might condition the diffusion of ideas about policy. Diffusion is likely to occur over various networks—in our case those linking experts and economists, businesses, and the environmental movement. These networks may convey contradictory information. The way information is combined at the domestic level is probably complex, and building a coalition for change may take time and involve forms of lobbying power not well captured by spatial models of legislative politics such as the one we employed here (Ward & Grundig, 2009). Current applications of spatial econometrics to policy diffusion proceed as if ideas travel only between governments, and they ignore the complexity and discontinuity of policy making. A potentially fruitful field for future research may be to try to capture some of this, using both process-tracing case studies and large-*N* methods.

The domestic factors we highlighted were special interest politics and the role of legislative medians. Although we found evidence that powerful energy lobbies constrain green taxation levels, future research could usefully focus on sectors that use energy intensively, which the measure we use does not capture. An interesting aspect of our results is that there is stronger evidence for the importance of legislative medians on the environmental dimension than on the left–right dimension in our preferred specification (Model 8). Future research could usefully be focused on other constitutional and institutional factors likely to affect the possibilities for changing policy, such as those found in presidential and strongly bicameral systems.

In regard to the international level, our most interesting finding is a negative one about tax competition: Countries with structurally similar trade patterns do not appear to influence each other. On the other hand, countries more tightly integrated into the globalized economy (as reflected by our measure *actual economic flows*) do have lower tax rates. Rather than competition with specific other countries (as captured by spatial lags), it seems likely that the perception of policy makers about a country's general competitive position matters. To disentangle these two effects, it may be necessary to look in detail at debates in key countries. We found significant evidence that policy ideas diffuse through international trade and IGO networks. Although this result parallels those in other articles in the diffusion literature, it is interesting that it holds up when other spatial effects are controlled for in an m-STAR estimation. Our study shows that, given the collinearity between spatial lags, it is dangerous to interpret results from single lag models as evidence for learning.

Green taxation is an important market-based mechanism for environmental regulation. Although it has long been suspected that the politics of special interests and legislative politics inhibit its use, leading to inefficiently low levels of taxation, ours is the first large-*N* study to provide rigorous evidence for these effects. Although the literature also suggests that states' policies affect each other, by using spatial lags we have been able to provide provisional estimates of how the effects of overcoming policy resistance in one country or group of countries might ramify through the system by spatial feedbacks. It is important that future efforts to project the impacts of introducing such measures as carbon pricing not only take account of interactions occurring through energy markets but also take account of the sorts of political interactions that we have shown to exist.

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Notes

1. We downloaded the data from http://www2.oecd.org/ecoinst/queries/TaxInfo.htm.
2. Imposing uniform standards fails to minimize the social cost of achieving any given target reduction because it fails to differentiate between those who can do much at little cost and those who cannot. Market-based regulation may also give polluters greater incentive to innovate.
3. Given the fact that there are potentially various domestic political variables and numerous ways by which they can condition the impact of international influences, the current article theorizes and tests only one particular conditional effect: whether the similarity in legislative median's ideological orientations between countries facilitates policy learning. We look at left–right and environmental emphasis dimensions of the legislative medians separately.
4. Holzinger, Knill, and Sommerer (2008) and Jahn (2009) include some measures that come under the green taxation heading in the policy indexes they construct, but these mainly weight institutional features and command and control regulation.
5. Because emulation and learning are difficult to distinguish in a large-*N* study, we use learning to cover both mechanisms.
6. Lobbies offer a schedule of bribes to a legislator, as a function of policy, the focus being on schedules revealing true preferences in the vicinity of equilibrium.
7. Binder and Neumayer (2005) find environmental pressure group strength cuts air pollution, allowing for potential endogeneity.
8. We consider only one aspect of veto power. For example, if agenda control is exerted by governments or committees, because they need not propose any changes, there is a second veto.
9. In theory, concerns about increased transboundary flows of "emission leakage" from competitor countries and about the global environment may complicate their calculations (Aidt, 2005; Conconi, 2003).
10. We follow standard notations for scalars (italic, e.g., w), vectors (bold lowercase, e.g., $\mathbf{w}$), and matrices (bold uppercase, e.g., $\mathbf{W}$).
11. Including a lagged dependent variable (LDV) generates simultaneity as the LDV is correlated with the error term; when country fixed effects are included, "Nickell bias" results. We do not first difference the data or use instruments for the LDV because we believe this bias is negligible given that we have time series of 11 years (Adolph, Butler, & Wilson, 2005) and because independent variable estimation often performs poorly in finite samples (Kiviet, 1995).

12. Real GDP is calculated by taking the product of real GDP per capita and population. Data on population are also drawn from the World Development Indicators. Real GDP per capita is in year 2000 dollars at purchasing power parity, based on Penn World Tables Version 6.2 (Heston, Summers, & Aten, 2006). A limitation of this measure is that it includes some forms of energy production that are not commonly subject to green taxation.
13. Results are available from the authors on request.
14. For the furthest left party, the lower end of the scale formed the left limit of the interval.
15. This procedure gives the mean position of members of the legislature, given our assumption about the spread of each party's members. It gives slightly different results from the alternative calculation of the median corrected for grouped data; but because the seat distributions are generally relatively symmetric, the two measures are highly correlated.
16. Spoon and Jensen (2009) use unweighted party position scores derived from the Comparative Manifestoes Project (CMP) in explaining divergence of EU member states from Kyoto goals, but this is to ignore the dominance within many coalitions of larger parties. Jahn (2009) also uses the CMP data when modeling his index of the institutionalization of environmental policy. Like Spoon and Jensen, he does not distinguish between majority and minority governments. Using arguments drawn from Tsebelis (2002) that the winset never decreases with the range of positions, he also includes the range of party positions. But the relationship between the range and policy change cannot be expected to be linear, nor even strictly monotonic, which makes the assumption of linear effects problematic.
17. These are (a) food and live animals directly for food; (b) beverages and tobacco; (c) crude materials, inedible, except fuels; (d) mineral fuels, lubricants, and related materials; (e) animal and vegetable oils, fats, and waxes; (f) chemical and related products; (g) manufactured goods, classified chiefly by material; (h) machinery and transport equipment; (i) miscellaneous manufactured articles; and (j) commodities and transactions not classified elsewhere. Data are from the UN Comtrade online database (United Nations, 2008).
18. Our decision to set negative values to zero affects the final connectivity matrix very little because there are very few values that are below zero.
19. We checked for influences through the general intergovernmental organization (IGO) network and found weaker effects.
20. We relied on Ingram, Robinson, and Busch's (2005) categorization to select environmental IGOs: Minimalist IGOs are those that have plenary meetings, committees, and possibly a secretariat, but without an extensive bureaucracy beyond research, planning, and information gathering.
21. Available at http://privatewww.essex.ac.uk/~ksg/data-5.html.

22. There are other more straightforward ways to capture the conditional effects of domestic variables on policy diffusion. For example, one can use the interaction between the domestic variable of interest (e.g., x) and the spatial lag that represents a diffusion effect of interest (**W**y). The estimation of this interaction term x***W**y could be easily implemented in the m-STAR model.
23. The OECD supplies free data at five yearly intervals, so we linearly interpolated, which is justifiable given there is little temporal variation in any particular country.
24. Fixed effects are often significant and large. For instance, for the United States, the value is –\$506.164 ($p < .01$), compared to the overall mean of \$614.56 for the dependent variable.
25. We found some slight evidence for nonlinearities when the effective median was far to the left, but these are not stable across specifications, so results are not reported here.
26. In Table 1 *green party* typically was not significant. Dropping it from the model in Table 2 makes little difference to the model reported in Table 2.
27. Although cap-and-trade emissions quotas have more predictable effects than pollution taxes given difficulties with estimating elasticities, green taxes raise revenue that can be used to subsidize environmental cleanup or to correct other distortions in the tax system; and they may be the only practical option with regard to citizen or consumer behavior (OECD, 2008a; Stavins, 2003).
28. The long-run equilibrium impact is given by $(\mathbf{I}_\mathbf{N} - \Sigma\rho_i \mathbf{W}_{\mathbf{it}} - \varphi\, \mathbf{I}_\mathbf{N})^{-1}\, \Delta\mathbf{x}_\mathbf{t}\, B$, where $\mathbf{I}_\mathbf{N}$ is the identity matrix, $\mathbf{W}_{\mathbf{it}}$ the submatrix of the i-th weighting matrix for period t, and $\Delta\mathbf{x}_\mathbf{t}\, B$ is the shock at time t. We assume the spatial weights and all other variables remain at 2003 values.

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