Mobility-based gerrymandering: Theory and evidence

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Abstract

This paper models theoretically and tests empirically the hypothesis that the locational choice of a "public bad" within a multi-tiered structure of government (a facility providing widespread benefits throughout the federation but inflicting damage to the locality hosting it) can be driven by strategic electoral considerations exploiting the heterogeneous migration responses to the location of the public bad by voters of different ideologies - a sort of mobility-based "gerrymandering." As long as the average utility loss from residing close to the public bad is larger for progressive voters than it is for conservative voters, conservative and progressive central governments will pursue opposite "cracking" strategies, the former locating the public bad in an electorally tight region to induce an exit of progressive voters and gain the region to the conservative party, the latter attempting to spread progressive voters out of safe and towards electorally tight regions. An application to waste treatment plant locations across Italian municipalities returns evidence in support of the model's main hypothesis.

Keywords—

JEL codes—

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1 Introduction

The original insight by Tiebout (1956) that decentralized provision of public goods encourages perfectly mobile households to reveal their preferences by moving to jurisdictions offering the desired mix of public services and associated tax prices (commonly referred to as "voting with their feet") has had a profound and lasting impact on both theoretical and empirical research in public and urban economics (Oates, 2006).

Among the innumerable refinements of the original set-up - most of which aimed at fostering its usefulness as a positive description of actual public good allocation mechanisms within decentralized government structures at the expense of its purely normative contribution - the ubiquitous presence of upper levels of government on top of the fragmented local one envisaged by Tiebout (1956) has prompted the discussion of the relevance of Tiebout-like migration within federal fiscal structures (Boadway and Tremblay, 2012, Agrawal et al., 2024).

In fact, the equilibrium allocation of households across a perfectly competitive arrangement of local jurisdictions in a federation will, in most circumstances, be systematically perturbed by the policy-making decisions of upper tiers of government. Of particular importance are the upper-tier of government's discrete decisions about the location of public facilities having heterogeneous spatial impacts through spillovers from the sites where the facilities are sited - frequently generating a mixture of a large-scale public good and a public bad having a modest spatial reach (e.g., nuclear waste repositories, power plants, incinerators, or airports). As a result, a key issue in that respect - and one that has attracted considerable academic interest - is bound to concern the decision-making process leading to the choice of the location of the spillover-generating facility and how the siting decision will deviate from the socially optimal one (Frey et al., 1996).

Conventional theoretical approaches to the facility siting issue have primarily focused on how local communities express their opposition. This is often seen as typical NIMBY (not-in-my-backyard) behavior. Less attention has been paid to how policymakers might anticipate the potential for residents to leave (exit decisions).

Many studies have applied the standard Coasian framework (Coase, 1960). In this approach, candidate communities negotiate directly for compensation in a market-like process without much involvement in political decision-making. Alternatively, some studies have relied on a framework of political competition among interest groups (Becker, 1983). Both approaches tend to reach the same efficient outcome: communities most affected by the public bad are the most motivated to oppose it, so the public bad is ultimately located in areas where residents are less concerned about its impact (Hamilton, 1993).

Later research has expanded these models by adding political economy factors, such as asymmetric lobbying, land ownership structures, and corruption. These studies have examined under what conditions political decisions about facility siting diverge from socially optimal ones (Fredriksson, 2000, Feinerman et al., 2004, Bellettini and Kempf, 2013).

In this paper, we take a different approach. We view the decision about the location of a facility generating a local public bad by an upper level of government in a federal structure as an instrument to influence the composition of the electorate in lower-level jurisdictions by exploiting the heterogeneous impact of the public bad on the utility of residents with different ideologies and their propensities to exit as a result - a sort of mobility-based "gerrymandering". In particular, we envisage a two-tier system of government (one central government and two regional governments) and two types of agents (progressives and conservatives) having an idiosyncratic attachment to each region and choosing in which of the two regions to live and consume one unit of housing. In addition, conservative and progressive agents are heterogenous with respect to income. A public bad (e.g., an incinerator) must be located in either of the two regions of the federation, with a noxious impact on the utility of residents of the region where it is located and no effect on the other region. As long as the average utility loss from residing close to the public bad is larger for progressive voters than it is for conservative voters, the location of the public good in a region will provoke a relatively larger outflow of progressive voters and will raise the share of conservative voters in that region. As a result, the model predicts opposite "cracking" strategies on the part of conservative and progressive central governments: a conservative-controlled central government will optimally locate the public bad in an electorally tight region to induce exit of progressive voters and help gain the region to the conservative party, while a progressive-controlled central government will locate it in a non-tight region in an attempt to costlessly (from an electoral point of view) spread progressive voters out towards tight regions, where their votes can turn out to be decisive. In addition, the model predicts a fall in property prices in the region hosting the public bad, while the impact on the level and distribution of income is uncertain.

Next, we test the model's implications on a rich dataset of new waste treatment plants in Italy (the 137 plant openings occurred during the years 2016-2019). While, as we discuss below, the theoretical insights of the model apply to a wide range of central siting decisions within a multilevel government structure, including both public goods and public bads having an impact on voters' location decisions, the opening of a waste treatment plant lends itself naturally to a test of the theory because of the following two features. First, in most circumstances - as the ones we deal with here - the decision is a top-down one. Despite the common provision of a formal participatory and decentralized decision process where involved communities and stakeholders can express their voice, the decision is ultimately in the hands of an upper level of government (the province, in the Italian case, as we discuss in section **3** below), so that electoral considerations from that level of government are likely to be a major issue at the final decision stage. Second, whether the waste treatment facility under consideration is a landfill, a waste-to-energy plant (an incinerator) or a biological conversion structure, the impact on the surrounding environment and the quality of life of the people living there in terms of pollution, smell, noise, congestion and aesthetics is bound to be devastating and irreversible, even in case the plant should be believed to possibly cease operating in the future. This implies that the eventual decision to relocate elsewhere, at a farther distance from the plant, is a plausible consequence of the siting decision.

The empirical analysis proceeds through the following three steps. First, we employ flexible non-parametric techniques to approximate the conditional mean function of (decade) long differences of election results, real estate prices, and income level and distribution along the spatial gradient originating from the new plant sites. Second, we parametrically quantify the impact of plant openings on "treated" and "spillover" localities, that is, on those localities that are close enough to plants to suffer a "direct" or "indirect" impact from them relative to control jurisdictions that are located at sufficiently far away distance not to suffer any consequence, but can still be assumed to be on parallel trends before the opening of the plants. The third and final step of the empirical analysis addresses the issue of whether the effects observed in the previous steps can be taken as physiological unintended consequences of random locations of new waste treatment plants across space or are driven by strategic considerations by the governments of the provinces (the upper level of government in charge of the decision of where new plants should be located), depending on their ideological affiliation.

The estimates show a steady decline in the share of the vote of the progressive party and property prices in the vicinity of new plant locations. Moreover, we find that the location of new plants is driven by heterogeneous forces depending on whether the province where the new waste management plant is to be located is ruled by a right-wing or by a left-wing government: right-wing controlled provinces locate the new plants in highly contested municipalities, while left-wing controlled provinces locate the new plants in sites that are close to municipalities with uncertain elections.

This paper relates to two distinct strands of literature. The first is the analysis of the partisan implementation of inefficient policies that are harmful to society but can be effective in "entrenching" an incumbent by preventing an adversary group from gaining power. Saint-Paul et al. (2016) address the question of under what conditions incumbent politicians might implement policies that damage their own natural constituencies. Within a traditional two-party framework (Left, Right) and rich and poor voters having different preferences over income redistribution, they show that, under some conditions, the Left may find it optimal to implement policies that, by reducing the income of their own constituency (the poor), make income redistribution even more valuable. However, this sort of political entrenchment differs from the mechanism we envision here in that it does not rely on cross-jurisdictional mobility of voters. Glaeser and Shleifer (2005) formalize the idea of an electorate-shaping strategy – labelled "the Curley effect" – consisting in differentially taxing different groups of voters to encourage emigration of one of the groups and maximize the share of the voters who support the incumbent – a policy that (as the ones implemented by four-times Boston's mayor James Michael Curley in the early 20^{th} century) successfully raises the incumbent's popularity while actually making both groups worse-off. Brueckner and Glazer (2008) consider instead the case where incumbents might intentionally adopt extremist policies hurting all groups of voters (such as over-provision of a public good), yet are particularly unattractive to a minority, leading some members of the minority to emigrate and thus reinforcing the political power of the initial majority for the next periods. While the underlying policy-induced migration mechanism in the presence of heterogeneous impacts of the policy in those papers is similar to ours, the distinguishing feature of our model is that it is the upper tier of government that chooses the location of a public bad to influence the electoral outcomes at the lower tier, both in the selected locality and in its surroundings.

The second strand of literature we contribute to is the voluminous empirical research on the socio-economic impact of unwanted facilities (or their closure) on local communities (Chay and Greenstone, 2005, Gleeson, 2007, Banzhaf and Walsh, 2008, Haninger et al., 2017, Mei, Gao, Zhang and Yang, 2021, Mei, Qiu, Wu and Meng, 2021, Rivera and Loveridge, 2022, Song et al., 2023). Besides estimating the consequences of siting plants on long differences of conventional economic variables (real estate prices and income level and distribution), we investigate for the first time the electoral impact of those siting decisions in terms of party shares at the municipal level at increasing distances from the new plant sites, both with non-parametric (Cattaneo et al., 2024, Butts, 2023*a*) and with parametric methods (Clarke, 2017, Butts, 2023*b*, Fiorini et al., 2024).

The rest of the paper is structured as follows. Section 2 develops the theoretical model and derives its main empirical predictions. Section 3 illustrates the data we use along with some descriptive spatial statistics, estimates the impact of the location of plants on a vector of long-differenced local outcomes, and tests the hypothesis of heterogeneous behavior of conservative versus progressive parties in plant location decisions. Section 4 discusses potential alternative theories of public bad siting in federations, and Section 5 concludes.

2 The Model

Our model first explores the impact of locating a public facility, such as a waste management plant, on the population distribution between two regions. We describe a two-tier system, with a central government and two regional governments. The population consists of two types of agents: *progressives* and *conservatives.* These agents are defined by their income levels, an idiosyncratic attachment to each region, and ideological views.

Each individual is characterized by an ideological view $i \in \{P, C\}$ describing a progressive and a conservative one. The population comprises N^P progressive individuals and N^C conservative ones, where $N^P + N^C = N$. The share of each type is $n^i = N^i/N$. Each individual earns an income w. Income distribution has a density $g^i(w)$. The wage distributions may differ between progressive and conservative individuals and lie on a support $[0, W^i]$. The average wage for agents with ideology i is $\bar{w}^i = \int_0^{W^i} wg^i(w)dw$ and the average wage in the total population is $\bar{w} = \sum_{i=P}^{C} n^i \bar{w}^i$.

Each individual lives in one of two regions $j \in \{1, 2\}$. An agent with ideology i living in Region j earns benefit β_j^i regardless of whether the facility is located in the region. This benefit is derived from the portfolio of public goods offered in the jurisdictions and other amenities. This match between ideology and a particular region can be referred to as the ideological location benefit.

Each individual also has personal preferences over the two regions. Regardless of ideology, each agent draws an idiosyncratic attachment index ε uniformly distributed on the support $[0, 1]$. The personal attachment index generates a benefit $\frac{1-\varepsilon}{2}$ of residing in Region 1 and $\frac{\varepsilon}{2}$ in Region 2.

A facility must be placed in one of the two regions. Define the facility index $F_j \in \{0, 1\}$, where 0 represents Region j having no facility and 1 for the cases where the region hosts it. Note that $F_2 = 1 - F_1$. We also use the notation $F_j = \emptyset$ to refer to the case where there is a complete absence of a facility. This will allow us to characterize the impact of the facility's arrival in one of the two regions. The facility impacts residents' utility in the region where it is placed. Denote by $V^i(F_j, w)$ the value for an agent of ideology i and wage w of living in a region j with the presence or absence of a facility. Denote by $\bar{V}^i(F_j) = \int_0^{W^h} V(F_j, w)g^i(w)dw$ the average benefit an individual with ideology i receives from living in a region with facility index F_j across all wages.

Each agent living in Region j consumes one housing unit out of a total of H_j units available. We assume that the aggregate housing supply is just enough to shelter the entire population i.e. $H_1 + H_2 = N$. We can then refer to the proportion of housing stock in Region j by $h_j = H_j/N$. The housing price in Region j as a function of facility index F_j is denoted by $Q_j(F_j)$.

The utility $U_j^i(F_j, w)$ for an agent of ideology $i \in \{P, C\}$ and income w living in Region $j \in \{1, 2\}$ with a facility index $F_j \in \{0, 1\}$ is an aggregation of various linear components described above. The utilities an agent of ideology i derives from a region's characteristics, net of housing price, are then given by

$$U_1^i(F_1, w) = \beta_1^i + V^i(F_1, w) - Q_1(F_1) + \tfrac{1-\varepsilon}{2} \qquad \text{and}$$
$$U_2^i(F_2, w) = \beta_2^i + V^i(F_2, w) - Q_2(F_2) + \tfrac{\varepsilon}{2}.$$

2.1 Demands for Location and Housing Market

Individuals prefer to live in the region that offers the highest utility. Ultimately, they weigh three attributes across the two regions. From now on, Region 1 is used as the reference point, and we express the location of the facility in terms of F_1 where $F_2 = 1 - F_1$. Individuals compare the ideological location benefits β_j^i. Denote by $\Delta\beta^i = \beta_1^i - \beta_2^i$ the relative attractiveness to region 1 for an agent with ideology i. A positive $\Delta\beta^i$ means that individuals with ideology i view Region 1 as a more desirable place to live, everything else being equal. The difference in housing prices between regions 1 and 2 for a given facility index F_1 is defined as $\Delta Q(F_1) = Q_1(F_1) - Q_2(1 - F_1)$. The difference in utility from living in a region with or without the presence of a facility for an individual with ideology i is represented by $\Delta\bar{V}^i(F_1) = \bar{V}^i(F_1) - \bar{V}^i(1 - F_1)$. If $\Delta\bar{V}^i(F_1)$ is positive, the facility benefits the nearby residents. A negative $\Delta\bar{V}^i(F_1)$ describes a case of a public bad, meaning that the value of living in a region with the facility is lower than residing away from it. Note that $\Delta\bar{V}^i(1) = -\Delta\bar{V}^i(0)$. In the complete absence of a facility, the difference in benefit from the facility is zero and is expressed by $\Delta\bar{V}^i(\emptyset) = 0$.

Lemma 1: *The housing demands $R_j^i\big(F_j; Q(F_1)\big)$ from individual with ideology $i \in \{P, C\}$ for living in Region $j \in \{1, 2\}$ with a facility index F_j, as function of housing price differential $Q(F_1)$ are*

$$R_1^i\big(F_1; \Delta Q(F_1)\big) = \frac{N^i}{2} + N^i\big[\Delta\bar{V}^i(F_1) + \Delta\beta^i - \Delta Q(F_1)\big], \tag{1}$$

$$R_2^i\big(F_1; \Delta Q(F_1)\big) = \frac{N^i}{2} - N^i\big[\Delta\bar{V}^i(F_1) + \Delta\beta^i - \Delta Q(F_1)\big]. \tag{2}$$

All proofs can be found in the appendix. As expected, an increase in $\Delta\beta^i$ means higher demand by individuals with ideology i for living in Region 1 to the detriment of Region 2. A reduction in price difference $\Delta Q(F_1)$ also stimulates housing demand in Region 1. Having the facility in a Region lowers its housing demand if $\Delta V^i(1)$ is negative. The demand increases when the facility provides a benefit instead of imposing harm. Note in the complete absence of a facility; housing demands are still given by (1) and (2), but where $\Delta\bar{V}^i(\emptyset) = 0$.

Each agent consumes one unit of the fixed housing supply H_j in one of the two regions. The market clearing condition in Region j is

$$R_j^P\big(F_j; \Delta Q^e(F_1)\big) + R_j^C\big(F_j; \Delta Q^e(F_1)\big) = H_j \tag{3}$$

Lemma 2: *The equilibrium price differential $\Delta Q^e(F_1)$ is given by*

$$\Delta Q^e(F_1) = \frac{h_2 - h_1}{2} + n^P[\Delta\bar{V}^P(F_1) + \Delta\beta^P] + n^C[\Delta\bar{V}^C(F_1) + \Delta\beta^C]. \tag{4}$$

An increase in housing supply in Region 1 lowers the equilibrium price differential. The housing supply in Region 2 has the opposite effect. Stronger preferences for Region 1 by individuals with ideology i (a higher $\Delta\beta^i$) lead to a higher equilibrium price differential. The facility's impact operates through $\Delta\bar{V}^i(F_1)$. If the facility is located in a given Region, the price differential diminishes when $\Delta\bar{V}^i(F_1)$ is negative. If the facility is beneficial, then the price differential will increase. In the complete absence of a facility, the equilibrium price differential is $\Delta Q^e(\emptyset) = \frac{h_2-h_1}{2} + n^P\Delta\beta^P + n^C\Delta\beta^C$.

Lemma 3: *The equilibrium share of progressive individuals living in Region 1 with facility index F_1 is given by*

$$\Gamma_1^P(F_1) = \frac{n^P(1-n^P)}{h_1}\left[\frac{h_1}{1-n^P} + [\Delta\beta^P - \Delta\beta^C] + [\Delta\bar{V}^P(F_1) - \Delta\bar{V}^C(F_1)]\right]. \tag{5}$$

Similarly, the equilibrium share of progressive individuals living in Region 2 is

$$\Gamma_2^P(F_1) = \frac{n^P(1-n^P)}{h_2}\left[\frac{h_2}{1-n^P} - [\Delta\beta^P - \Delta\beta^C] - [\Delta\bar{V}^P(F_1) - \Delta\bar{V}^C(F_1)]\right]. \tag{6}$$

Obviously, $\Gamma_j^C(F_1) = 1 - \Gamma_j^P(F_1)$.

A higher proportion of progressive individuals in the total population (n^P) leads to an increase in the proportion of individuals with that political view in both regions. The same relationship applies to the proportion of conservative individuals. The term $\Delta\beta^P - \Delta\beta^C$

dictates when individuals with a particular political view prefer to live when the facility's location is removed from the equation. A higher value of $\Delta\beta^P - \Delta\beta^C$ means that progressive individuals are relatively more attracted to Region 1 than conservative ones. This leads to an increase in the proportion of progressive individuals living in Region 1. Assumption 1 below guarantees a positive share of residents with a given political view in each region in the complete absence of a facility.[1] The term $\Delta\bar{V}^P(F_1) - \Delta\bar{V}^C(F_1)$ is particularly important, as it describes the impact of placing the facility in a given region on the distribution of each type across both regions. We look at such impact in the following subsection. Assumption 1 also guarantees a positive share of residents with a given political view in each region with the presence of a facility.[2] This implies that the relative impact of the facility on residents' welfare cannot be too extreme.

Assumption 1: We assume that $[\Delta\beta^P - \Delta\beta^C] + [\Delta\bar{V}^P(F_1) - \Delta\bar{V}^C(F_1)]$ is bounded by $\min\left\{\frac{-h_1}{1-n^P}, \frac{-h_2}{n^P}\right\}$ and $\max\left\{\frac{h_1}{n^P}, \frac{h_2}{1-n^P}\right\}$.

2.2 Regional Impacts of the Facility

Since our empirical work compares long differences, it is important to conceptualize our static model to make the proper comparison. To discuss the impact of the localization of the facility on different variables, we compare those variables for the case where there is no facility at all to the case in which the facility is located in one or the other region. We know that in the complete absence of a facility, both terms $\Delta\bar{V}^i(\emptyset)$ equal zero. We also know that the housing equilibrium price differential is $\Delta Q^e(\emptyset) = \frac{h_2 - h_1}{2} + n^P\Delta\beta^P + n^C\Delta\beta^C$ in such a case. Only proportions and preferences β^i_j determine the equilibrium price and the allocation of individuals across the two regions.

What happens when the facility is placed in a region depends on how the two types of individuals are affected. Imagine a simple case when the facility provides a positive benefit to progressive individuals $\Delta\bar{V}^P(1) > 0$ and no benefit to conservative ones $\Delta\bar{V}^C(1) = 0$. Locating the facility in Region 1 leads to a higher proportion of progressive individuals in the region. Progressive individuals are willing to suffer from increased housing prices because they benefit from the facility. The conservative ones do not, so they want to move away. The proportion of progressive individuals in Region 1 will increase, as well as the relative housing price in favour of Region 1.

We will now describe the particular case we investigate in our empirical analysis. Assumption 1 below describes a case where everyone is harmed by having the facility in their

[1]See the Proof of Lemma 3 where $F_1 = \emptyset$.

[2]See conditions (18) and (19)in Proof of Lemma 3.

region, but progressive individuals suffer more.

Condition 1: Imagine that $\Delta\bar{V}^P(1) < \Delta\bar{V}^C(1) < 1$.

Proposition 1: *Given Assumption 1, the equilibrium housing price differential $\Delta Q^e(F_1)$ is negatively affected by the localization of the facility in Region 1 and positively affected by the localization in Region 2.*

Because the facility is undesirable to all residents where it is placed, the arrival of a facility reduces the residential demand in the host region. This will put upward pressure on the housing price in the facility-free region. Price differential $\Delta Q^e(F_1)$ decreases when the facility is placed in Region 1 and increases when Region 2 is the recipient. Note that only the fact that the facility causes harm drives the result in Proposition 1. No matter the relative harms faced by progressive and conservative individuals, this result holds.

Proposition 2: *Given Assumption 1, the region receiving the facility sees a reduction in the share of progressive individuals $\Gamma_j^P(F_1)$. In contrast, the other region experiences an increase in the share of such individuals.*

The trade-off faced by residents of the region who receive the facility is as follows: by leaving, an individual can avoid suffering the harm caused by the facility but will have to pay a higher housing cost. Progressive individuals who suffer more from the facility are generally more willing to accept the higher housing cost. People with progressive views are willing to pay more to live away from the facility, creating an exodus of such individuals from the region with the facility. At the same time, the share of conservative individuals increases in the hosting region. Individuals with conservative views care less about the impact of the facility, so they are more sensitive to housing prices. Such individuals move to the region with the facility to enjoy lower housing costs. Obviously, the region spared from the facility experiences the exact opposite effects.
We can now look at the impact of the location of the facility on average wages.

Proposition 3: *The impact of locating a facility in Region j on the average income in Region 1 is*

$$\begin{aligned} E_1(w|F_1) - E_1(w|\emptyset) = \; & \textstyle\sum_{i=P}^{C} \; n^i \int_0^W \Delta V^i(F_1, w) w g^i(w) dw \\ & -\Big[n^P \Delta\bar{V}^P(F_1) + n^C \Delta\bar{V}^C(F_1)\Big]\bar{w}. \end{aligned} \tag{7}$$

The impact on income of locating the facility in a given region is ambiguous. To understand the forces at work, imagine that $V^i(F_1, w)$ is independent of wages. The impact of the facility on average wages, $\sum_{i=P}^{C} V^i(F_1)n^i\bar{w}^i - [n^P\Delta\bar{V}^P(F_1) + n^C\Delta\bar{V}^C(F_1)]\bar{w}$, is still ambiguous, unless the average wages across a political view are the same. In such a case, all individuals of a political view make the same location decision no matter their wage. Still, the proportion of each type of individual and the associate average wages will influence average regional wages. This is a composition effect across political views. For example, if progressive individuals are wealthier and represent a larger share of the population, then placing the facility in Region 1 will lead to an increase in the average wage in Region 2 and a reduction of the same variable in Region 1. When $V^i(F_1, w)$ depends on wages, individuals of a given political view make location decisions according to their wage. This is an in-group composition effect. If wealthier conservative individuals tend to suffer more from the facility, most individuals who are toward the facility have low income, contributing to a reduction in the average wage in the receiving region.

2.3 Political Strategies Using Facility Location

We will now look at the political incentives associated with the placement of a facility. We will show that the political allegiance of the upper-tier government responsible for the facility can influence its placement choice. To do so, we introduce a simple two-tier electoral process. It starts with an incumbent government with ideology i representing the upper-tier jurisdiction in charge of locating the facility. At some point after the facility is placed, regional elections take place. Before such elections, some individuals move according to their preferences. In each region, a progressive and a conservative politician are running for office, with one of the two parties holding office. Residents in a region vote according to their political views and some other random factor, represented by some white noise process. A candidate with a progressive agenda wins the regional election if $\Gamma_j^P(F_1) + \psi[\Psi - 1/2] \geq \frac{1}{2}$, where Ψ follows a two-two Beta distribution. Like a normal distribution, the two-two Beta has a concave density distribution function, given by $6\Psi(1-\Psi)$, but is bounded on the support zero to one. Its cumulative distribution function is $3\Psi^2 - 2\Psi^3$. The term $\psi[\Psi - 1/2]$ represents a white noise with a zero average and where small deviations from the fundamental are more likely than larger ones.[3] The parameter ψ describes the importance of the noize. The candidate with a conservative agenda then wins if $\Gamma_j^C(F_1) - \psi[\Psi - 1/2] > \frac{1}{2}$.

We assume that the upper-tier government only cares about maximizing the expected number of regional elected officials sharing the same political view. The voting shares of the last regional elections are known and given by $\Gamma_j^i(\emptyset)$. Assuming that the upper-tier

[3] If the white noise followed a uniform distribution (zero-zero Beta distribution), placement of the facility would have no impact on the expected number of regional election won because of the linearity it generates.

government knows voters' preferences over the facility, it can accurately anticipate their relocation choices.

To understand some of the electoral strategic forces at play, we start with a simple example where efficient placing of the facility does not matter and there is no voting noise ($\psi = 0$). Suppose that in the last electoral round the progressive candidate in Region 1 won by twenty-five percentage points and the progressive candidate in Region 2 lost by one percentage point. Consider a facility that provides a benefit for progressive voters and no benefit for conservative ones. In such a case, a progressive upper-tier government could ensure that both progressive candidates win by placing the facility in Region 2. As long as Region 2 attracts progressive voters by at least two percentage points and displaces conservative voters to Region 1 by no more than twenty-five percentage points, both elected officials will have a progressive view. A conservative government may not be able to ensure two wins at the regional level, but placing the facility in Region 1 can solidify the win in Region 2. Consequently, upper-tier governments with different political views make different placement decisions to induce voter movement and have a regional electoral gain.

A complete characterization of facility placements would not be tractable as it depends on vote shares and the impact of the facility on voter movements. There would simply be too many cases to consider. We concentrate on the case described by Assumption 1 to help shed light on the empirical results that follow.

Lemma 4: *The expected number of progressive regional officials is* $EM^P(F_1) = \frac{\frac{\psi-1}{2}+\Gamma_1^P(F_1)}{\psi} + \frac{\frac{\psi-1}{2}+\Gamma_2^P(F_1)}{\psi}$ *and the equivalent expectation for conservative officials is* $EM^C(F_1) = \frac{\frac{\psi-1}{2}+\Gamma_1^C(F_1)}{\psi} + \frac{\frac{\psi-1}{2}+\Gamma_2^C(F_1)}{\psi}$.

By locating the facility in a given Region, the upper-tier government changes the relative ideological composition in both regions. Consequently, an upper-tier government can use the location of the facility for electoral purposes.

Proposition 4: *A progressive upper-tier government wishing to maximize the expected number of regional elections won by progressive candidates will place the facility in the region with the highest* $\frac{\left[1-2\Gamma_j^P(\emptyset)\right]^2-\psi^2}{h_j}$. *A conservative upper-tier government with a similar objective will place the facility in the region with the lowest* $\frac{\left[1-2\Gamma_j^C(\emptyset)\right]^2-\psi^2}{h_j}$.

We start by looking at the case where both regions have the same housing stock. A progressive government would then place the facility in the region with the largest $\left[1-2\Gamma_j^P(\emptyset)\right]^2$, i.e., the region with the widest vote share difference in the last election. By locating the facility in a region with a large majority, the progressive government induces progressive voters to move to the region with a tight majority to secure it or to switch it in its favor. A conservative government follows exactly the opposite strategy. It will place the facility in the region with the lowest vote share difference to induce exit of progressive voters and gain the region to the conservative candidate.

Finally, the size of the region matters because a given group of individuals moving to a region will have a bigger impact on the voting share the smaller the destination region is. This is why the placement condition is divided by h_j. For a progressive government, placing the facility in a small region is desirable because you need all you can get to move the voting share in the larger region where it sends the progressive voter. The exact opposite is true for the conservative government, which wants to bring voters to a large region with a tight majority. As long as the size differences is not that large, voting shares drive the results.

3 Empirical analysis

We test our model on data from the openings of waste treatment plants in Italy during the past decade. The set-up lends itself to a test of this kind because of the Italian multi-level structure of government and of the top-down and irreversible nature of the siting decision that is bound to have severe consequences on the welfare of the people living in the vicinity of the designated site.

Italy is divided into around 100 provinces that, according to the environmental law 152 of 2006, are responsible for selecting the sites for the location of waste management plants (art. 197-1-d).[4] In addition, provincial elections (direct popular election of the head of the province) have traditionally been characterized by an extremely high degree of ideological content, there have never been instances of non-partisan races, and, in spite of the fact that Italy has a multi-party system and coalition governments are frequent, all elected provincial governments can univocally be identified as left-wing or right-wing.[5] As a result, we can test whether the location of new plants is driven by heterogeneous forces depending on whether the province where the new waste management plant is to be located is ruled by a right-wing or by a left-wing government.

[4] According to the same law, the upper level of government - the 20 regions - is in charge of defining the general principles (related to local geomorphological structure, prevalent land use, or presence of historic, cultural and environmental heritage) for the identification of the areas that are *not* suitable for the location of new waste management plants.

[5] The provincial structure and electoral rules subsequently underwent radical changes (creation of special metropolitan authorities and introduction of indirect election of the head of the province) in 2014, after the crucial decision-making period that we consider here.

3.1 Exploratory analysis: nonparametric binned regressions

We view the opening of new waste treatment plants as the outcomes of political-economic processes leading to the realization of rare events at specific spatial locations (local political-administrative jurisdictions - the Italian municipalities in our case) and generating two distinct types of effects on surrounding jurisdictions, depending on the distance of those localities from the plant sites. First, we expect the closer localities to be *directly* affected by the start of the operation of the plant (generation of noise, smell, pollution, congestion), with residents in those areas suffering a utility loss and possibly deciding to relocate as a result.[6] Second, we expect farther away localities to be affected *indirectly* by the opening of the plant, in the sense that they do not suffer any direct consequence from the plant itself (the above noxious consequences are likely to dissipate within a short distance from the plants), but only from the perturbations to their labor, housing, and political-electoral markets that are provoked by the relocation of the people that were living close to the plants and chose to move.[7] Finally, we postulate the validity of SUTVA (stable unit treatment value assumption), that is, the existence of a cluster of municipalities that are located at a long enough distance from the plant so that the relocation of households from the areas close to the plant is likely to be so diluted that no effect will display there, yet share substantial similarities with the treated and spillover localities in terms of underlying socio-economic environment. This is conceptually similar to the general equilibrium approach of Minton and Mulligan (2024) to the estimation of difference-in-differences models with market-based spillovers, where contamination of the within-market control group is addressed via selection of an additional, outside-market control group not subject to spillovers yet experiencing common pre-treatment trends.

3.1.1 Binscatters

We observe whether a new waste treatment facility opens during a given time interval $(\underline{t}, \overline{t})$ in locality $i = 1, ..., M$ (M being the country's total number of municipalities) and code that information in a binary variable $p_i \in \{0, 1\}$. The analysis uses the $\overline{M} < M$ observations on municipalities m that are located at distance $d_{m,i} < \overline{d}$ from localities i where $p_i = 1$. Since relying on cross-sectional variation in outcomes between municipalities that happen

[6] This amounts to a familiar (negative) technological externality.

[7] In practice, from a structural point of view, we can rule out neither that the localities in the directly treated cluster suffer indirect effects too due to the neighborhood turnover process, nor that the noxious effects of the plants have a direct impact on some of the localities in the spillover cluster. However, our empirical analysis aims at the reduced-form estimates of the effects of plant openings on a vector of relevant socio-economic variables in the two clusters (rings, as we define below) that are selected in a data-driven way.

or not to have a new plant in the surroundings would suffer from likely endogeneity issues due to unobserved differences between treated and non-treated areas in the presence of non-random choice of the site of a new plant, we focus on the relatively small areas surrounding a site that is selected for a new plant and exploit the within-area variation in outcomes from before to after the event at increasing distances from the site (Linden and Rockoff, 2008, Hornbeck and Keniston, 2017, Muehlenbachs et al., 2015, Currie et al., 2015, Diao et al., 2017, Haninger et al., 2017, Bartik et al., 2019, Mei, Qiu, Wu and Meng, 2021, Shr et al., 2023).

We rely on a long-differences approach (Chay and Greenstone, 2005, Gleeson, 2007, Banzhaf and Walsh, 2008) in the sense that we construct the pre-post (log) difference $\Delta \ln(y_m) = \ln(y_{m,\bar{t}+k_1}) - \ln(y_{m,\underline{t}-k_2})$ for a number of outcome variables y and observe their trajectories along a distance gradient originating in the plant location. In particular, we make an exploratory analysis through binscatter plots as non-parametric approximations of the conditional mean function $E\left(\Delta \ln(y_m) | d_{m,i}\right)$.

To do so, letting $d_{(j)}$ denote the j^{th} order statistic of the sample $\left(d_{1,i}, d_{2,i}, ..., d_{\overline{M},i}\right)$ and $\lfloor \cdot \rfloor$ denote the floor operator, we partition the support of $d_{m,i}$ into J bins $\widehat{B}_1, \widehat{B}_2, ..., \widehat{B}_J$, based on the marginal empirical quantiles of $d_{m,i}$:

$$\widehat{B}_j = \begin{cases} \left[d_{(1)}, d_{(\lfloor M/J \rfloor)}\right) & j = 1 \\ \left[d_{(\lfloor M(j-1)/J \rfloor)}, d_{(\lfloor Mj/J \rfloor)}\right) \quad \text{if} & j = 2, ..., J-1 \\ \left[d_{(\lfloor M(J-1)/J \rfloor)}, d_{(M)}\right) & j = J \end{cases} \tag{8}$$

and for each bin $\widehat{B}_j$ we compute:

$$\begin{aligned} \widehat{\delta}_j &= E\left(\Delta y_m \middle| d_{m,i} \in \widehat{B}_j\right) \\ &= \frac{1}{M_j} \sum_{m=1}^{M} \mathbf{1}\left(\Delta y_m \in \widehat{B}_j\right) \Delta y_m \end{aligned} \tag{9}$$

This amounts to partitioning-based least squares estimation of a k-degree polynomial within each bin (a $k = 0$ degree spline, or a piecewise constant fit, in this case). The resulting scatter plot of means $\widehat{\delta}_j$ shows how the effects of the new plants (the change in the variable y_m from before to after the opening of the waste management facilities) vary with distance from them.

Indeed, the choice of J poses the familiar bias-variance trade-off in nonparametric estimation: a sufficiently large J is required to attenuate the approximation bias arising from the hypothesis of constant within-bins fit, but variance increases with J due to falling bin-specific sample size. Cattaneo et al. (2024) propose to select the J that minimizes the integrated mean square error (IMSE) in a data-driven way:

$$\widehat{J}_{IMSE} = \left\lceil \left(\frac{2\mathcal{B}_M}{\mathcal{V}_M} \right)^{\frac{1}{3}} M^{\frac{1}{3}} \right\rceil \tag{10}$$

where the terms $\mathcal{V}_M$ and $\mathcal{B}_M$ denote the asymptotic variance and (squared) bias of the binscatter, respectively. If $J < \widehat{J}_{IMSE}$, the binscatter will show less variability at the expense of a larger bias (undersmoothing), and vice-versa .

3.1.2 Data

The key information that we exploit is the opening of waste management plants in Italy during the period 2016-2019. There were 137 openings of new plants during that period (Figure 1), while the existing number of plants at the beginning of the period (year 2015) was 442 (Figure 2).

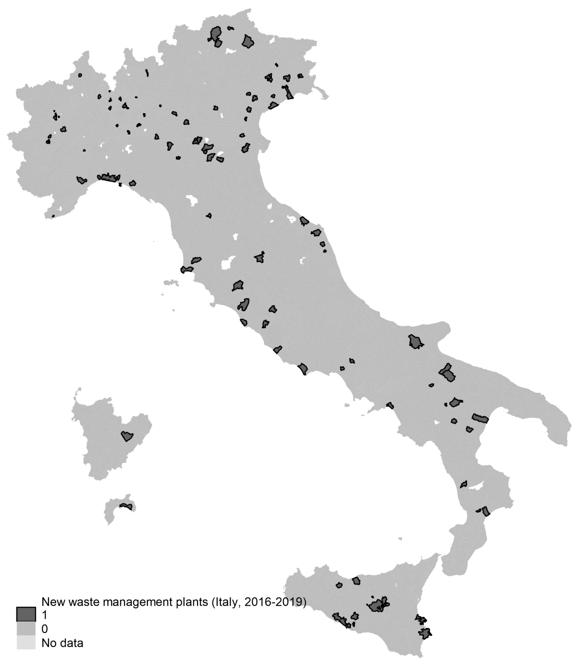

Figure 1: New plants (2016-2019)

Figure 1 and Figure 2 show that both the initial stock of waste management plants and the location of new plants is rather evenly spread out across the entire national territory.

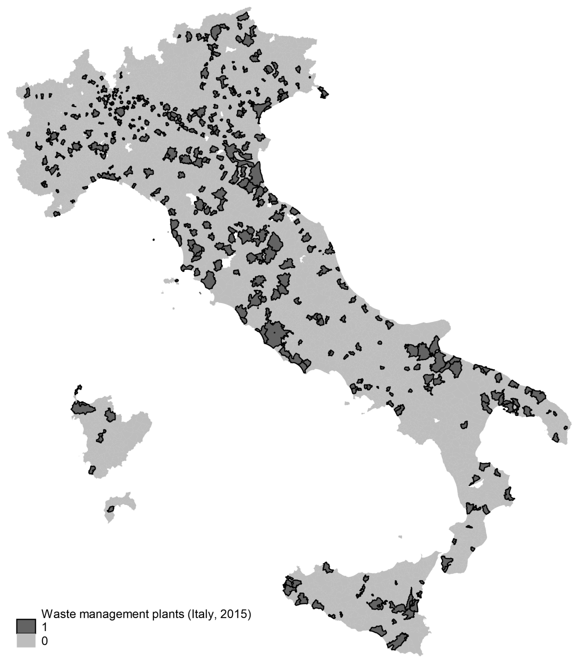

Figure 2: Existing plants (2015)

Waste management plants belong to either of the following three types: 1) sanitary landfills or monitored dumpsites; 2) waste-to-energy plants producing electricity, heat, or transport fuels via thermal conversion or incineration; 3) biological conversion plants generating biogas or compost. All of those plants, though to a varying degree, have a negative impact on the surrounding environment and on the quality of life of the people living there in terms of pollution, smell, noise, congestion, and aesthetics. Due to the relatively small number of observations on new plant openings, we treat all plant types alike and group them into a single binary variable.

As for the impact on surrounding localities, we use data on: a) electoral results b) real estate prices; c) personal income level and distribution. These are discussed in turn below.

a) electoral results

First, to verify if the location of plants provokes a change in the ideological composition of the population, we use the change in the share of the vote of the progressive (centre-left) coalition between the parliamentary elections of 2008 and the parliamentary elections of 2022, both measured at the municipal level. This is the closest proxy of the local ideological affiliation of the electorate given that most mayoral elections are non-partisan, so that they provide little information on the actual ideological composition of the population.

b) housing prices

We observe average prices of real estate at the municipal level for both domestic and business properties. Information on real estate values comes from the database provided by the Italian Internal Revenue Authority (http://www.agenziaentrate.gov.it). For each municipality, a minimum-maximum range of market values is provided (price per square meter) for each type of property and state of maintenance and conservation. We rely on the long difference Δ(2011-2021) and focus on the average percentage change in the values of domestic and business properties (Δlog).

c) personal income level and distribution

Income data for tax purposes at the municipal level - yearly taxable income per taxpayer = gross labor and capital income minus tax deductions and allowances - are available from the Department of Finance of the Italian Treasury (http://www1.finanze.gov.it). For each municipality, we observe the value of income and the number of taxpayers with income falling within each of the tax brackets set by the national government: up to €15,000 yearly gross income; €15,000 to €28,000; €28,000 to €55,000; €55,000 to €75,000; above €75,000. We rely on the long difference Δ(2011-2021) and focus on percent change (Δlog) in average income per capita and percent change (Δlog) in the number of taxpayers with income above €55,000.[8]

Descriptive statistics of the variables used in the analysis are in Table 1, where we report the average gross percentage change in the outcome variables from before the opening of plants (variable years from 2008-2011, depending on data availability, as discussed above) to after the opening of plants (years 2021-2022), at the distance intervals from the new 137 plant sites (the rings) around which the spatial differencing analysis of Section 3.2 is centered.

Table 1 documents some well-known changes that the Italian economy and society experienced in the recent decades (Bloise et al., 2024), including general upward trends in income and downward trends in real estate prices and progressive party share of the vote. We remove those common trends by taking municipal-level deviations from sample mean long-differences for all variables.

[8]We use the €55,000 thresholds instead of the top income threshold because the latter would produce a large number of zeros in smaller localities where no taxpayers with those income levels reside.

Table 1: Descriptive statistics: long differences (2011-2021)

	0-4km	4-8km	8-30km
Δshare(left vote)	-0.114	-0.098	-0.097
	(0.073)	(0.069)	(0.069)
Δln(domestic property price)	-0.143	-0.105	-0.122
	(0.159)	(0.150)	(0.157)
Δln(business property price)	-0.190	-0.152	-0.154
	(0.213)	(0.176)	(0.179)
Δln(income p.c.)	0.055	0.059	0.061
	(0.051)	(0.047)	(0.053)
Δln(taxpayers$>$ €55k)	0.283	0.328	0.302
	(0.194)	(0.203)	(0.231)

Notes: left vote share: 2008-2022 (parliamentary election results - progressive coalition - at the municipal level).

3.1.3 Results

The binscatter regressions are reported in Figure 3 to Figure 7. The number of bins in each graph is the IMSE-minimizing one and vertical segments represent 95% confidence intervals (Cattaneo et al., 2024). Remarkably, all binscatters exhibit a common pattern of: 1) a *direct* effect on the outcomes of interest reaching up to about 3 to 5 km from the sites where the new plants open; 2) a *spillover* effect of the opposite sign (whose size and significance shows some physiological variability across the different outcome variables considered) in the area at about up to 6 to 10 km distance from the plants; 3) the virtual absence, for all outcome variables, of any further effect at longer distances from the sites. The fact that the direct effects of the plant openings have a spatial reach of no more than a few kilometres is in line with recent estimates of the housing price gradient from policies of remediation and revitalization of contaminated sites and conversion and energy transition of polluting plants (Currie et al., 2015, Haninger et al., 2017, Mei, Gao, Zhang and Yang, 2021, Rivera and Loveridge, 2022, Song et al., 2023). The evidence that those direct effects appear to be generally reversed in the immediately outer region is compatible with an hypothesis of proximate relocation of households as a result of the opening of the plant.

Figures 3 to 7 First, the decade-long trajectory of the share of the parliamentary vote of the progressive coalition (Figure 3) falls by 2% in the close vicinity of the plants (around

0-3 km), recovers slightly at around 5 to 10 km distance, and shows no significant deviation from the common trend at longer distances.

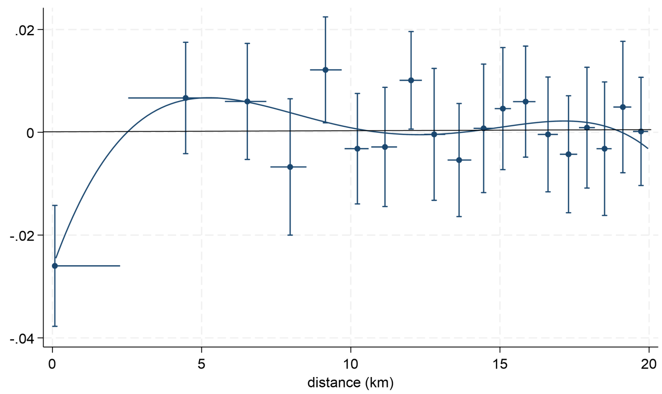

Figure 3: Progressive party vote (share)

As for the consequences on the real estate market, the binscatters provide a stark picture of falling prices in the areas immediately surrounding the plant sites (0-3 km, around -4% to -5%) and rising prices in the spillover area of 3 to 6 km (around +2% to +3%) both for domestic (Figure 4) and business (Figure 5) properties.

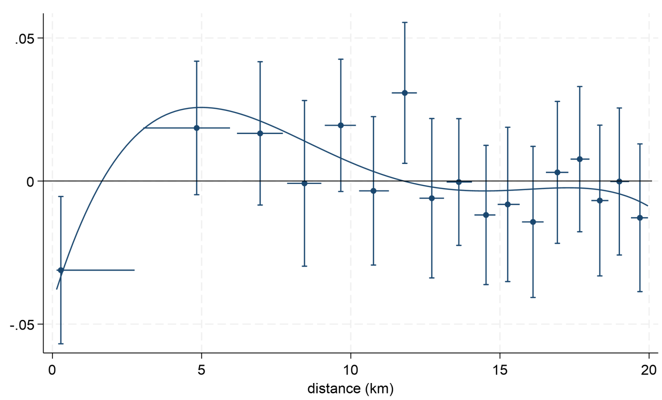

Figure 4: Domestic property price

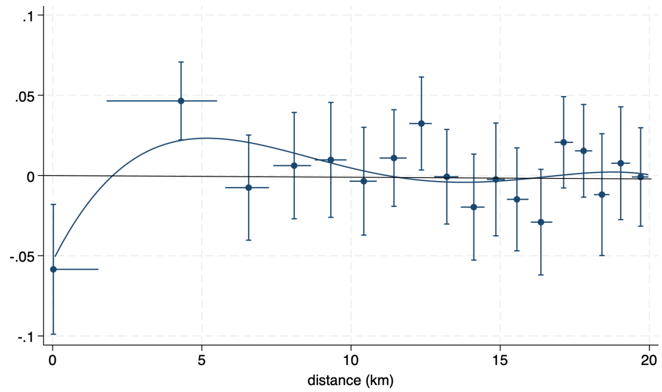

Figure 5: Business property price

These real estate price effects are compatible with the hypothesis of an amenity-driven turnover of the local population, with the opening of the plant pushing out residents of a higher economic status to farther away unaffected areas. Those social groups' higher demand for building upgrading, renovation, and improvement raises property prices there (an "environmental gentrification" spillover from the opening of a plant), the opposite ("reverse gentrification") taking place in the localities that are closer to the plants (Banzhaf and McCormick, 2007). In addition, the average impoverishment of the communities in the vicinity of the new plants induces a change in the composition of tastes for private consumption goods too. In turn, this is reflected into the local retail and service sector, as witnessed by the fall in business property values (a reverse "Starbucks effect": O'Sullivan (2005)). No effects on the real estate market can be discerned at longer distances from the plants.

Finally, growth of income per capita across the decade 2011-2021 (Figure 6) shows a slight fall (less than 1%) in the vicinity of the plant (0 to around 6 km), some increase in the outer spillover region (about 6 to 10 km), and a flat pattern thereafter, though the mean binscatter long-differences are not significantly different from the baseline zero trend. A similar pattern emerges as far as the trajectory of the number of high-income taxpayers (Figure 7) is concerned, that falls by about 3% in the proximity of the plant (0 to 3 km) and shows some tendency to increase - by around 2 to 4% - in the spillover area (6-9 km).

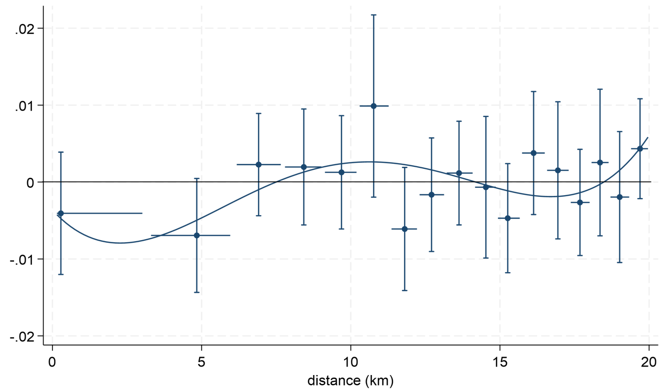

Figure 6: Per capita income

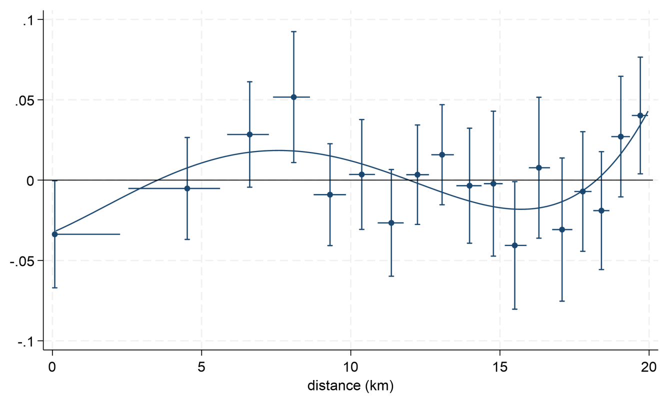

Figure 7: High-income taxpayers (number)

Overall, the binscatter evidence is consistent with the hypothesis that the opening of waste treatment plants provokes a relocation of left-wing oriented and relatively well-off individuals from areas close to new plant sites to areas that are just far enough from new plants not to suffer their negative consequences.

3.2 Spatial Difference-in-Differences: the rings method

The exploratory analysis in the previous section suggests that the opening of new plants generates: i) *direct* effects on the communities at close distance from the new plant sites; ii) *indirect* effects of the opposite sign, likely due to households' proximate resettlements, on farther away communities; iii) virtually no visible effect at longer distances from the plants. This makes it possible to parameterize the exposure of communities to the effects from the new plant openings through properly designed "rings" (treatment, spillover, control) within a spatial differencing approach (Clarke, 2017, Butts, 2023*b*, Fiorini et al., 2024).[9] In particular, the nonparametric analysis suggests a *buffer* ring with a radius d^T of about 3 to 5 km, a *spillover* ring with a radius d^S of up to 6 to 10 km from plants, and a *control* ring including observations at farther away distances that can be considered unaffected by the plant openings (yet on parallel trends with the directly and indirectly affected localities before the plant openings, a hypothesis that we formally test in the Appendix).

As a result, we model the impact of the opening of a plant on the long difference of the outcome as in equation (11):

$$\Delta \ln(y_m) = \gamma + \delta^T D\left(d_{m,i} \leq d^T\right) + \delta^S D\left(d^T < d_{m,i} \leq d^S\right) + \varepsilon_m \tag{11}$$

with δ^T and δ^S capturing the average treatment and spillover effects in the two inner rings.[10]

The estimation results are reported in Table 2, where we experiment with three radius lengths (3, 4, 5 km for the treated area; 6, 8, 10 km for the spillover area). The control group region includes municipalities located outside the spillover area and at up to 30 km

[9] Alternative approaches to accounting for potential spatial spillovers on control observations include Kline and Moretti (2014), Delgado and Florax (2015), and Greenaway-McGrevy and Phillips (2023). In their analysis of the long run effects of a major US regional development program - the Tennessee Valley Authority (TVA) - Kline and Moretti (2014) compare the economic performance of counties within the geographical scope of the TVA (treated units) to that of outside counties (control units), and drop all counties that border the TVA region because of possible spillovers from the program. This strategy is not suitable to our paper, where estimation of the sign and size of the spillover is central to the analysis. Second, Delgado and Florax (2015) allow for spatial interaction in the outcomes of treated and untreated units within a 'local' contiguity area that does not extend to the entire spatial system. The fact that SUTVA holds outside that area allows estimation of average direct and indirect effects of treatment depending on the proportion of treated neighbors. In our paper, while we do not exclude the existence of endogenous interactions in local outcomes between adjacent localities, we model instead the reduced form effects of the shocks to those outcomes (the environmental impact of the opening of new plants) as a function of distance from the origins of those shocks under the assumption of dissipation (validity of SUTVA) beyond a threshold distance. Finally, in their study of the impact of upzoning on housing construction in Auckland, New Zealand - a set-up where spillovers from upzoned areas ($\frac{3}{4}$ of residential land) to non-upzoned areas show no tendency to dissipate with distance - Greenaway-McGrevy and Phillips (2023) employ a partial identification strategy that uses pre-treatment trends in the control group (non-upzoned areas) to extrapolate a set of counterfactual outcomes that are used to bound the magnitude of the spillover effect. In the environment that we study, though, the evidence of highly localized spillovers allows construction of proper control groups where SUTVA holds.

[10] Since the hedonic price function describing the equilibrium market-clearing relationship between real estate prices and the local public good (the quality of the environment) is likely to change over time due to the neighborhood turnover generated by the shock to the local public good (Banzhaf and McCormick, 2007) - an issue that, in our context, arises almost by design and that is exacerbated by our use of long differences - our reduced-form estimates of real estate price responses to the opening of new waste treatment plants can only recover a pure capitalization effect that cannot be given a welfare (marginal willingness to pay) interpretation (Kuminoff and Pope, 2014, Banzhaf, 2001).

from plants, for a total number of observations of around 5,000.[11] The standard errors are clustered by the regions surrounding each of the new plants.

Table 2: Spatial DiD estimates

	δ^T	δ^S	δ^T	δ^S	δ^T	δ^S
	(0-3 km)	(3-6 km)	(0-4 km)	(4-8 km)	(0-5 km)	(5-10 km)
left vote (share)	-0.024***	0.001	-0.017***	-0.001	-0.012***	0.001
	(0.005)	(0.006)	(0.005)	(0.005)	(0.004)	(0.004)
property (dom.)	-0.031***	0.021	-0.021*	0.017	-0.008	0.010
	(0.012)	(0.014)	(0.012)	(0.013)	(0.010)	(0.012)
property (bus.)	-0.046***	0.020	-0.036**	0.002	-0.023*	0.001
	(0.017)	(0.016)	(0.016)	(0.014)	(0.013)	(0.012)
income (p.c)	-0.006*	-0.008**	-0.006*	-0.002	-0.009***	-0.000
	(0.004)	(0.004)	(0.004)	(0.003)	(0.003)	(0.003)
rich (number)	-0.026*	-0.002	-0.019	0.026**	-0.012	0.022**
	(0.016)	(0.013)	(0.015)	(0.013)	(0.012)	(0.011)

Notes: Control group $<$ 30 km. Standard errors clustered by plant site. *** : p value $<$ 0.01; ** : p value $<$ 0.05; * : p value $<$ 0.10.

As expected, the estimates of δ^T and δ^S closely reflect the patterns of the binscatters in figures 3 to 7, with the performance of the three rings specifications differing depending on the outcome variables considered. As for property prices and left vote share, the estimated direct effects fall in size and/or significance as the treated region δ^T grows, compatibly with the hypothesis that the observations that receive the most intense treatment dose (the ones that are closest to the plants) dilute into samples being subject to spillover forces going in the opposite direction.

Taking the smallest inner treatment ring as a reference (0-3 km), the vote share of progressive parties falls by $2\frac{1}{2}\%$, while domestic and business property prices fall by 3% and almost 5% respectively relative to control localities. The estimates of the effects on income are in the expected direction (falling per capita income and share of high-income taxpayers), though they turn out to be less precisely estimated. The spillover effects tend to have

[11] The results are almost unchanged when using more (up to 50 km) or less (20 km) extensive control group regions.

the opposite signs as the direct effects, but they are generally not statistically significant, plausibly due to the more dispersed destination patterns of the outmoving population.

3.3 Are plants located strategically?

Finally, we investigate whether electoral considerations drive province-level decisions about where to locate new plants. In particular, as long as the environmental degradation following the location of new waste treatment plants provokes a relatively higher exit rate of left-wing voters than right-wing ones, strategic electorate-shaping use of plant locations to maximize the chances of winning municipal elections requires right-wing controlled provinces to locate the new plants in highly contested municipalities to provoke outmigration of left-wing voters and help gain those municipalities to the right. On the other hand, left-wing controlled provinces should locate the new plants in sites that are close to municipalities with uncertain elections to induce left-wing voters to move to those localities and improve the chances of left-wing mayors of winning the elections there.

To test this "cracking" hypothesis, we estimate a cross-sectional binary dependent variable model that makes the choice of the location of a plant in a given locality depend on a vector of exogenous local characteristics and electoral and political variables measured *before* the decision about the location of the new plants (the previous decade). In particular, we focus on two political economy variables that (as highlighted by the theoretical model) can be believed to affect the location of a plant for strategic purposes. The first is the electoral uncertainty characterizing municipal elections, proxied by the average win margin of the mayor (standardized vote difference between the elected mayor and the most voted opponent, ranging from 0 - in case of an exact tie - to 100 - in case of all votes going to the winning candidate or of uncontested races) in the elections occurred during the 2000-2010 decade. This should capture the underlying structural uncertainty of local elections before the decision about the location of the plants is made. The second variable is the ideology of the provincial government (progressive/conservative) at the last elections held for provincial councils in the second half of the 2000s decade.[12]

The results are presented in Table 3 for the whole sample and in Table 4 and Table 5 for right-controlled and left-controlled provinces respectively. Each table includes a specification that controls for the local degree of uncertainty of elections only (columns (1) and (2)), a specification that allows for a spillovers from uncertainty of elections in the municipalities located up to a 10 km distance (columns (3) and (4)) and a specification that also allows for a further away *placebo* spillover of uncertainty of elections in the municipalities located more

[12] Due to the existence of a staggered five-years length electoral calendar, provincial elections took place between 2007 and 2011.

than 10 km and up to a 50 km distance (columns (5) and (6)). Along with the electoral variables, we include a number of local socio-economic and demographic characteristics that might contribute to driving the location decisions. In all tables, odd-numbered columns report linear probability model estimates, and even-numbered columns report Probit model estimates (marginal probability effects).

In general (Table 3), new plants are more likely to be located in municipalities that have larger population size, lower population density, and a lower proportion of employment in agriculture, while the actual production of waste and residents' socio-economic indicators are not estimated to play any significant role. The results in Table 4 and Table 5, though, show that the political-economic process leading to the choice of the location of new plants is systematically different in right-wing versus left-wing controlled provinces. In the former (Table 4), plants are significantly more likely to be located in municipalities characterized by high electoral uncertainty. The effect of electoral uncertainty in nearby municipalities is negligible. On the other hand, in left-wing controlled provinces (Table 5), plants are significantly more likely to be located in municipalities that are surrounded by municipalities with uncertain elections.

The evidence is compatible with a hypothesis of strategic use of plant locations to shape the electorate so as to maximize the chances of winning municipal elections. Right-wing controlled provinces tend to locate new plants in highly contested municipalities to provoke outmigration of left-wing voters and gain those municipalities to the right. Left-wing controlled provinces tend to locate new plants in sites that are close to municipalities with uncertain elections to induce left-wing voters to move out to those localities and improve the chances of left-wing mayors of winning the elections there.

Table 3: Plant location: all provinces

	(1)	(2)	(3)	(4)	(5)	(6)
	LPM	Probit	LPM	Probit	LPM	Probit
population	0.001	0.001**	0.001	0.001**	0.001	0.001**
	(0.001)	(0.000)	(0.001)	(0.000)	(0.001)	(0.000)
population density	-0.001*	-0.001	-0.001*	-0.001	-0.001*	-0.001
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
agricolture (% empl.)	-0.033***	-0.056***	-0.031***	-0.056***	-0.031***	-0.055***
	(0.007)	(0.017)	(0.007)	(0.017)	(0.008)	(0.017)
foreign (%)	0.005	0.006	0.005	0.006	0.005	0.006
	(0.006)	(0.005)	(0.006)	(0.005)	(0.006)	(0.005)
unemployment (%)	0.008	0.016	0.002	0.009	-0.001	0.005
	(0.013)	(0.015)	(0.013)	(0.015)	(0.014)	(0.015)
waste p.c.	0.002	0.003	0.002	0.003	0.002	0.003
	(0.006)	(0.005)	(0.006)	(0.005)	(0.006)	(0.005)
vote gap	-0.004	-0.006	-0.001	-0.003	-0.001	-0.002
	(0.004)	(0.004)	(0.003)	(0.004)	(0.003)	(0.004)
vote gap (0-10km)			-0.022**	-0.028**	-0.020*	-0.025**
			(0.010)	(0.011)	(0.012)	(0.012)
vote gap (10-50km)					-0.011	-0.019
					(0.021)	(0.023)
observations	6,539	6,539	6,539	6,539	6,539	6,539

Notes: Dependent variable: discrete indicator of the opening of a new plant in a municipality during the years 2016-2019. Cross-sectional estimates on the sample of municipalities not hosting waste treatment plants in 2015. Standard errors clustered by region. *** : p value < 0.01; ** : p value < 0.05; * : p value < 0.10.

Table 4: Plant location: right-wing provinces

	(1)	(2)	(3)	(4)	(5)	(6)
	LPM	Probit	LPM	Probit	LPM	Probit
population	0.001	0.001**	0.001	0.001***	0.001	0.001***
	(0.001)	(0.000)	(0.001)	(0.000)	(0.001)	(0.000)
population density	-0.001*	-0.001	-0.001*	-0.001*	-0.001*	-0.001
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
agricolture (% empl.)	-0.030***	-0.060***	-0.028***	-0.061***	-0.029***	-0.061***
	(0.008)	(0.019)	(0.008)	(0.019)	(0.007)	(0.019)
foreign (%)	0.006	0.006	0.006	0.007	0.006	0.007
	(0.007)	(0.006)	(0.007)	(0.007)	(0.008)	(0.007)
unemployment (%)	0.011	0.020	0.007	0.015	0.007	0.015
	(0.016)	(0.016)	(0.015)	(0.015)	(0.016)	(0.015)
waste p.c.	0.003	0.003	0.003	0.002	0.003	0.003
	(0.006)	(0.006)	(0.006)	(0.006)	(0.006)	(0.006)
vote gap	-0.010**	-0.016**	-0.008**	-0.013**	-0.008*	-0.013*
	(0.004)	(0.007)	(0.004)	(0.007)	(0.004)	(0.007)
vote gap (0-10km)			-0.014	-0.021	-0.015	-0.020
			(0.013)	(0.016)	(0.015)	(0.018)
vote gap (10-50km)					0.004	-0.004
					(0.024)	(0.027)
observations	4,246	4,246	4,246	4,246	4,246	4,246

Notes: Dependent variable: discrete indicator of the opening of a new plant in a municipality during the years 2016-2019. Cross-sectional estimates on the sample of municipalities not hosting waste treatment plants in 2015. Standard errors clustered by region. *** : p value < 0.01; ** : p value < 0.05; * : p value < 0.10.

Table 5: Plant location: left-wing provinces

	(1)	(2)	(3)	(4)	(5)	(6)
	LPM	Probit	LPM	Probit	LPM	Probit
population	0.001	0.001	0.001	0.001	0.001	0.001*
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.000)
population density	0.001	0.001	0.001	0.001	0.001	0.001
	(0.002)	(0.001)	(0.002)	(0.001)	(0.002)	(0.001)
agricolture (% empl.)	-0.029	-0.050	-0.031	-0.052	-0.034	-0.055
	(0.027)	(0.039)	(0.027)	(0.039)	(0.029)	(0.040)
foreign (%)	0.007	0.009	0.006	0.009	0.005	0.007
	(0.013)	(0.013)	(0.013)	(0.013)	(0.012)	(0.012)
unemployment (%)	0.009	0.016	-0.004	-0.001	-0.027	-0.029
	(0.029)	(0.032)	(0.029)	(0.033)	(0.034)	(0.042)
waste p.c.	-0.006	-0.005	-0.007	-0.006	-0.006	-0.005
	(0.006)	(0.008)	(0.006)	(0.008)	(0.005)	(0.008)
vote gap	0.006	0.004	0.010	0.010	0.015**	0.015**
	(0.007)	(0.008)	(0.006)	(0.007)	(0.006)	(0.006)
vote gap (0-10km)			-0.035***	-0.043***	-0.027**	-0.033**
			(0.011)	(0.016)	(0.012)	(0.015)
vote gap (10-50km)					-0.082*	-0.112*
					(0.043)	(0.059)
observations	2,293	2,293	2,293	2,293	2,293	2,293

Notes: Dependent variable: discrete indicator of the opening of a new plant in a municipality during the years 2016-2019. Cross-sectional estimates on the sample of municipalities not hosting waste treatment plants in 2015. Standard errors clustered by region. *** : p value < 0.01; ** : p value < 0.05; * : p value < 0.10.

4 Discussion

As briefly mentioned in the introduction, the existing literature has tackled the issue of the location of unwanted facilities by means of models that virtually ignore the potential *exit* of residents in the vicinity of selected sites and focus instead on their *voice* responses. In this section, we consider a potential alternative political economy explanation of the process leading to the choice of the location of a public bad in a federation that shares a number of similarities with our theoretical set-up (the theory of special interests politics), test its empirical predictions on our data, and compare them to the ones that derive from our mobility-based gerrymandering model.

The popular theory of special interests (pork barrel) politics studying the determinants of the adoption of tactical distributive policies (projects) that concentrate benefits at a small scale while spreading costs across the rest of the polity (Shepsle and Weingast, 1981) typically assumes competition between two parties differing from each other in their issue positions (ideology) and their redistributive promises to the voters. The latter are in turn distinguished by the degree of their ideological affinity to one party or the other (Left/Right) and care heterogeneously about particularistic benefits that can temper their basic party loyalties (Dixit and Londregan, 1996). Within that framework, politicians from either party will either target their core constituents by directing to them the benefits of redistributive policies ("take care of their own") or take loyal supporters' vote for granted and direct instead redistributive benefits to groups of "swing voters". Within a multi-level structure of government as the one we consider here, the theory predicts that central legislators will opt for a biased distribution of centralized resources across localities with the sole purpose of gaining an electoral advantage for local officials of their own side, such as the approval of projects whose benefits are concentrated in a given district, thus boosting the popularity of the incumbent there, while the project costs will be spread out to all of the other districts (Johansson, 2003).

Indeed, while the distributive policy we discuss here implies inflicting a damage rather than providing a flow of benefits to the community that is selected to host the facility (and generating spread-out benefits rather than burdens to the rest of the country), yet the specularity of our environment to the pork barrel politics one makes the hypothesis that the siting decision be driven by maximization of local electoral support (minimization of local electoral loss) on the part of the upper-tier decision-maker a plausible one. When applied to the public bad location choice, pork barrel politics theory leads, under reasonable assumptions, to the following two empirical predictions.

First, if the central decision-maker having to make a one-off decision about the location of a public bad in one of the regions of the federation across which a population of ideological voters is unevenly spread is risk-averse and is only concerned about “locking in” its immobile core constituents (Cox and McCubbins, 1986), then it should select the region with the smallest share of its own voters. Importantly, this strategy ought to be exactly the same for right-wing and left-wing central governments, and it implies siting the unwanted facility in regions with a large majority of the other party - a sort of “machine politics” pattern(Dixit and Londregan, 1996) that is entirely different from the one deriving from our model.

Second, one should observe that the siting decision produces different consequences depending on the ideology of the local incumbent: the share of the vote of the progressive party should fall only in jurisdictions that are controlled by progressive local governments, because some of the progressive voters will blame the local incumbent for getting the noxious facility and switch to the other party or to abstention; in jurisdictions controlled by a conservative local government, the share of progressive voters should instead increase because of a similar behavior on the part of conservative voters, thus leading to different predictions than our model does.

To test the first empirical implication, we re-estimated the discrete choice model while controlling, besides the uncertainty of local elections, for the municipal-level share of the progressive party in the parliamentary elections of 2008. Neither in conservative-controlled nor in progressive-controlled provinces, the location of the plant appears to be affected by the municipal-level share of the progressive party. As shown in Table 6, the coefficient is not significantly different from zero in any of the specifications, and its inclusion has no consequences on the estimation of either of the other parameters, including the electoral uncertainty ones that are confirmed to be the main drivers of provincial plant siting decisions and, as predicted by our model, have heterogeneous effects on the location of plants depending on the ideology of the provincial government.

Next, to verify if the consequences of siting the plants are heterogeneous depending on the control of the municipality, we re-estimated Equation 11 separately on two subsamples. Though the prevalence of non-partisan mayoral elections prevents us from directly observing whether mayors are supported by conservative or progressive majorities in the city councils, we can still (though imperfectly) infer their ideology by using the results of national elections at the municipal level, where, unlike mayoral elections, party affiliation reveals itself in a clear way. Based on the parliamentary elections occurred in 2008, before the opening of the new plants, and splitting the sample into progressive-majority municipalities and conservative-majority ones, the estimated impact on the share of the vote of the progressive party (long difference between the 2008 and the 2022 elections) in the immediate vicinity of the plant - parameter δ^T (0-3 km) - equals -0.018 (standard error $= 0.007$) in the former and -0.013 (standard error $= 0.005$) in the latter, thus lending little support to the second prediction of the pork barrel politics model.

Table 6: Plant location: local ideology

	(1)	(2)	(3)	(4)	(5)	(6)
	right-wing provinces			left-wing provinces		
population	0.001	0.001	0.001	0.001	0.001	0.001
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
population density	-0.001*	-0.001*	-0.001*	0.001	0.001	0.001
	(0.001)	(0.001)	(0.001)	(0.002)	(0.002)	(0.002)
agricolture (% empl.)	-0.030***	-0.028***	-0.028***	-0.028	-0.030	-0.032
	(0.009)	(0.009)	(0.008)	(0.028)	(0.028)	(0.029)
foreign (%)	0.006	0.007	0.007	0.006	0.005	0.004
	(0.007)	(0.007)	(0.007)	(0.011)	(0.012)	(0.011)
unemployment (%)	0.007	0.004	0.004	0.007	-0.006	-0.028
	(0.016)	(0.015)	(0.017)	(0.027)	(0.027)	(0.033)
waste p.c.	0.004	0.003	0.003	-0.006	-0.007	-0.006
	(0.007)	(0.007)	(0.007)	(0.006)	(0.006)	(0.005)
progressive vote share	0.022	0.021	0.021	0.014	0.012	0.012
	(0.016)	(0.017)	(0.017)	(0.018)	(0.017)	(0.016)
vote gap	-0.010**	-0.008*	-0.008*	0.006	0.010	0.014**
	(0.004)	(0.004)	(0.004)	(0.007)	(0.006)	(0.006)
vote gap (0-10km)		-0.013	-0.014		-0.035***	-0.027**
		(0.013)	(0.015)		(0.011)	(0.012)
vote gap (10-50km)			0.003			-0.082*
			(0.023)			(0.044)
observations	4,244	4,244	4,244	2,291	2,291	2,291

Notes: Dependent variable: discrete indicator of the opening of a new plant in a municipality during the years 2016-2019. Cross-sectional estimates on the sample of municipalities not hosting waste treatment plants in 2015. Standard errors clustered by region. *** : p value $<$ 0.01; ** : p value $<$ 0.05; * : p value $<$ 0.10.

5 Conclusions

This paper has modelled theoretically and tested empirically the hypothesis that, within a multi-tiered structure of government, the upper level of government's decision about the location of a facility generating a local public bad be driven by the desire to modify the composition of the electorate in lower-level jurisdictions to favor ideologically aligned local administrators. In particular, the mechanism rests on the hypothesis that the decision-maker at the upper tier exploits the heterogeneous impact of the public bad on the utility of residents with different ideologies and their propensities to exit as a result - a sort of mobility-based "gerrymandering".

The model predicts that, as long as the average utility loss from residing close to the public bad is larger for progressive voters than it is for conservative voters, opposite "cracking" strategies will be adopted by conservative and progressive central governments: a conservative-controlled central government will optimally locate the public bad in an electorally tight region to induce an exit of progressive voters and gain the region to the conservative party, while a progressive-controlled central government will locate it in a non-tight region in an attempt to costlessly (from an electoral point of view) spread progressive voters out towards tight regions, where their votes can turn out to be decisive.

The empirical analysis on new waste treatment plant openings in Italian municipalities points to a fall in the share of the vote of the progressive party and in property prices in the immediate vicinity (3-5 km) of new plant locations, a spillover effect of the opposite sign in the slightly farther away area (about up to 6 to 10 km distance from the plants), and the virtual absence of any effect at longer distances from the sites, compatibly with the hypothesis of relocation of left-wing oriented and relatively well-off individuals from areas close to new plant sites to areas that are just far enough from new plants not to suffer their negative consequences. Moreover, we find that the location of new plants is driven by heterogeneous forces depending on whether the province where the new waste management plant is to be located is ruled by a right-wing or by a left-wing government, with right-wing controlled provinces locating the new plants in highly contested municipalities, and left-wing controlled provinces locating the new plants in sites that are close to municipalities with uncertain elections. The implications from alternative theories of public bad siting in federations receive no empirical support.

Finally, it should be stressed that, while lending themselves naturally to the analysis of the siting of a public bad such as a polluting facility, the theoretical insights of our paper apply to a far larger set of non-environmental issues that, by heterogeneously influencing the location decisions of mobile households of different ideologies, could be strategically maneuvered for electoral purposes. This includes controversial top-down decisions in multilevel structures of government such as the allocation of asylum seekers to local communities, the siting of divisive places of worship, or the adoption of desegregation educational policies. While these policy choices have indeed attracted considerable attention in the recent political economy literature (Hainmueller and Hopkins, 2014, Dustmann et al., 2019, Dazey and Gay, 2024), approaching them through the lens of our mobility-based gerrymandering model promises to contribute to our comprehension of their genesis and consequences.

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Appendix: theoretical model

Proof of Lemma 1: First note that $F_2 = 1 - F_1$ and consequently $\Delta\bar{V}^i(1-F_1) = -\Delta\bar{V}^i(F_1)$. An agent of ideology i and wage w is indifferent to living in either region if and only if

$$U_1^i(F_1, w) + \frac{1-\varepsilon}{2} = U_2^i(1-F_1, w) + \frac{\varepsilon}{2}. \tag{12}$$

For each wage w and indeology i, we can define

$$\hat{\varepsilon}^i(w) = \frac{1}{2} + \Delta V^i(F_1, w) + \Delta\beta^i - \Delta Q(F_1). \tag{13}$$

An agent with wage w and ideology i prefers living in Region 1 if and only if $\epsilon \leq \hat{\varepsilon}^i(w)$. Using the defintion of $\Delta\bar{V}^i(F_1)$ we have that

$$\int_0^W \hat{\varepsilon}^i(w) g^i(w) dw = \frac{1}{2} + \Delta\bar{V}^i(F_1) + \Delta\beta^i - \Delta Q(F_1). \tag{14}$$

Therefore $R_1^i\big(F_1; Q(F_1)\big) = N^i \int_0^W \hat{\varepsilon}^i(w) g^i(w) dw$ is given by (1) and $R_2^i\big(F_1; Q(F_1))\big) = N^i \int_0^W [1-\hat{\varepsilon}^i(w)] g^i(w) dw$ is given by (2). ■

Proof of Lemma 2: The market clearing conditions in Regions 1 and 2 are described by the two following equations where $\Delta Q(F_1) = Q_1(F_1) - Q_2(1-F_1)$

$$\Delta Q(F_1) = \frac{1}{2} - \frac{H_1}{N} + \frac{N^P}{N}[\Delta\bar{V}^P(F_1) + \Delta\beta^P] + \frac{N^C}{N}[\Delta\bar{V}^C(F_1) + \Delta\beta^C], \tag{15}$$

$$\Delta Q(F_1) = \frac{H_2}{N} - \frac{1}{2} + \frac{N^P}{N}[\Delta\bar{V}^P(F_1) + \Delta\beta^P] + \frac{N^C}{N}[\Delta\bar{V}^C(F_1) + \Delta\beta^C]. \tag{16}$$

The right-hand side of the two equations above are equal. We can obtain expression by adding the two equations (4). ■

Proof of Lemma 3: Evaluated at $\Delta Q^e(F_1)$, housing demands divided by N are

$$\frac{R_1^i(\cdot)}{N} = n^i h_1 + n^i \Big[\Delta\bar{V}^i(F_1) + \Delta\beta^i - n^P[\Delta\bar{V}^P(F_1) + \Delta\beta^P] - n^C[\Delta\bar{V}^C(F_1) + \Delta\beta^C]\Big],$$

$$\frac{R_2^i(\cdot)}{N} = n^i h_2 - n^i\Big[\Delta\bar{V}^i(F_1) + \Delta\beta^i - n^P[\Delta\bar{V}^P(F_1) + \Delta\beta^P] - n^C[\Delta\bar{V}^C(F_1) + \Delta\beta^C]\Big].$$

The equilibrium share of progressive individuals living in Region 1 for a given F_1 is

$$\Gamma_1^P(F_1) = \frac{R_1^P\big(F_1; \Delta Q^e(F_1)\big)}{R_1^P\big(F_1; Q^e(F_1)\big) + R_1^C\big(F_1; \Delta Q^e(F_1)\big)}. \tag{17}$$

Using housing demands evaluated at $\Delta Q^e(F_1)$, we can show that $\Gamma_1^P(F_1)$ is given by expression (5). We solve for $\Gamma_2^P(F_1)$ in the same way, where $F_2 = 1 - F_1$ and $\Delta\bar{V}^i(1-F_1) = -\Delta\bar{V}^i(F_1)$. Finally, $\Gamma_j^C(F_j) = 1 - \Gamma_j^P(F_j)$. We can show that $\Gamma_1^P(F_1) \in]0,1[$ iif

$$-\frac{h_1}{1-n^P} - [\Delta\beta^P - \Delta\beta^C] < [\Delta\bar{V}^P(F_1) - \Delta\bar{V}^C(F_1)] < \frac{h_1}{n^P} - [\Delta\beta^P - \Delta\beta^C]. \tag{18}$$

Similarly, we can show that $\Gamma_2^P(F_1) \in]0,1[$ iif

$$-\frac{h_2}{n^P} - [\Delta\beta^P - \Delta\beta^C] < [\Delta\bar{V}^P(F_1) - \Delta\bar{V}^C(F_1)] < \frac{h_2}{1-n^P} - [\Delta\beta^P - \Delta\beta^C]. \tag{19}$$

■

Proof of Proposition 1: Using the definition of $Q^e(F_1)$ can then show that $Q^e(0) > Q^e(\emptyset) > Q^e(1)$. ■

Proof of Proposition 2: In absence of a facility, $\Gamma_1^P(\emptyset)$ is given by equation (5) where $\Delta\bar{V}^P(1) - \Delta\bar{V}^C(1) = 0$. When the facility is placed in Region 1, $\Gamma_1^P(1)$ decreases because $\Delta\bar{V}^P(1) - \Delta\bar{V}^C(1) < 0$. At the same time, we can see from (6) that $\Gamma_2^P(1)$ increases. ■

Proof of Proposition 3: The average wage in Region 1 for a given facility location index F_1 when evaluated at $\Delta Q^e(F_1)$ is $E_1(w|F_1) = \sum_{i=P}^{C} n^i \int_0^W \hat{\varepsilon}^i(w) w g^i(w) dw$, where

$$E_1(w|F_1) \quad = \sum_{i=P}^{C} n^i \int_0^W \Big[\tfrac{1}{2} + \Delta\beta^i + \Delta V^i(F_1, w) - \Delta Q^e(F_1)\Big] w g^i(w) dw.$$

$$\begin{aligned} E_1(w|F_1) = \quad & \sum_{i=P}^{C} n^i\Big[h_1 + \Delta\beta^i - n^P[\Delta\bar{V}^P(F_1) + \Delta\beta^P] - n^C[\Delta\bar{V}^C(F_1) + \Delta\beta^C]\Big]\bar{w}^i \\ & + \sum_{i=P}^{C} n^i \int_0^W \Delta V^i(F_1, w) w g^i(w) dw. \end{aligned} \tag{20}$$

We can also show that in the complete absence of a facility, the average wage in Region 1 is

$$E_1(w|\emptyset) = \sum_{i=P}^{C} n^i\Big[h_1 + \Delta\beta^i - n^P\Delta\beta^P - n^C\Delta\beta^C\Big]\bar{w}^i. \tag{21}$$

Using the definition of $\bar{w}$, the impact of locating the facility in one of the regions is given by equation (7). ■

Proof of Lemma 4: The expected number of elections won by a progressive official EM^P is given by

$$\begin{aligned} EM^P = \quad & \left[Prob\left(\Psi \geq 1/2 + \tfrac{1/2-\Gamma_1^P(F_1)}{\psi}\right)\left(1 - Prob\left(\Psi \geq 1/2 + \tfrac{1/2-\Gamma_2^P(F_1)}{\psi}\right)\right)\right] 1 \\ & + \left[\left(1 - Prob\left(\Psi \geq 1/2 + \tfrac{1/2-\Gamma_1^P(F_1)}{\psi}\right)\right) Prob\left(\Psi \geq 1/2 + \tfrac{1/2-\Gamma_2^P(F_1)}{\psi}\right)\right] 1 \\ & + \left[Prob\left(\Psi \geq 1/2 + \tfrac{1/2-\Gamma_1^P(F_1)}{\psi}\right) Prob\left(\Psi \geq 1/2 + \tfrac{1/2-\Gamma_2^P(F_1)}{\psi}\right)\right] 2 \end{aligned} \tag{22}$$

The expected number of progressive officials is

$$\begin{aligned} EM^P(F_1) = \quad & 2 - \tfrac{[1+\psi-2\Gamma_1^P(F_1)]^2}{4\psi^3}[2\psi - 1 + 2\Gamma_1^P(F_1)] \\ & - \tfrac{[1+\psi-2\Gamma_2^P(F_1)]^2}{4\psi^3}[2\psi - 1 + 2\Gamma_2^P(F_1)]. \end{aligned} \tag{23}$$

Similarly, the expected number of conservative officials is give by

$$\begin{aligned} EM^C(F_1) = \quad & \tfrac{[1+\psi+2\Gamma_1^C(F_1)]^2}{4\psi^3}[2\psi - 1 - 2\Gamma_1^C(F_1)] \\ & + \tfrac{[1+\psi+2\Gamma_2^P(F_1)]^2}{4\psi^3}[2\psi - 1 - 2\Gamma_2^C(F_1)]. \end{aligned} \tag{24}$$

The vote shares for both parties are positive as long as $\Gamma_j^P(F_1) \in]\psi/2, 1-\psi/2[$. A minor variation on Assumption 1 would gurantee positive vote shares. ■

Proof of Proposition 4: We start by assesing the impact of $[\Delta\bar{V}^P(F_1) - \Delta\bar{V}^C(F_1)]$ on EM^P. First note that

$$\frac{\partial EM^P}{\partial \Gamma_j^P(F_1)} = \frac{3}{2\psi^3}\left[\psi^2 - \left[1 - 2\Gamma_j^P(F_1)\right]^2\right]. \tag{25}$$

Consequently, $\frac{\partial EM^P}{\partial[\Delta\bar{V}^P(F_1) - \Delta\bar{V}^C(F_1)]}$ is given by the expression bellow.

$$\frac{3n^P(1-n^P)}{2\psi^3}\left[\frac{\psi^2 - \left[1 - 2\Gamma_1^P(F_1)\right]^2}{h_1} - \frac{\psi^2 - \left[1 - 2\Gamma_2^P(F_1)\right]^2}{h_2}\right]. \tag{26}$$

We know that $[\Delta\bar{V}^P(\emptyset) - \Delta\bar{V}^C(\emptyset)] = 0$ and $[\Delta\bar{V}^P(1) - \Delta\bar{V}^C(1)] < 0$. Placing the facility in region 1 leads to a increase in the expected number of elections won by a progressive candidate if an only if

$$\frac{\left[1-2\Gamma_1^P(\emptyset)\right]^2-\psi^2}{h_1} > \frac{\left[1-2\Gamma_2^P(\emptyset)\right]^2-\psi^2}{h_2}. \tag{27}$$

The exact opposite is true for placing the facility in Region 2. Consequently, placing the facility in the region with the highest $\frac{\left[1-2\Gamma_j^P(\emptyset)\right]^2-\psi^2}{h_j}$ maximizes EM^P. Similarly, we can show that a conservative government wish placing the facility in the region with the lowest $\frac{\left[1-2\Gamma_j^P(\emptyset)\right]^2-\psi^2}{h_j}$ ■

Appendix: Parallel trends

Denote by $y_{mt}(\tau_m, \sigma_m(\tau))$ the potential outcome of municipality m at time $t \in \{0,1\}$ as a function of municipality m's own treatment-status $\tau_m \in \{0,1\}$ (with $\tau_m = 1$ if $d_{m,i} \leq d^T$) and of the entire vector of treatment assignments in the spatial structure, where $\tau \in \{0,1\}^M$ denotes the M-dimensional vector of all municipalities' treatments and $\sigma_m(\cdot)$ captures the exposure of m to that vector of treatments (the spillover). Treatment occurs between periods 0 and 1. Then, assume the following:

Assumption 1. *No spillover effect on treated units ($\tau_m = 1$):*

$$y_{mt}(1, \sigma_m(\tau)) = y_{mt}(1, \sigma_m(\mathbf{0})) \tag{28}$$

Assumption 2. *SUTVA (stable unit treatment value assumption): existence of untreated units ($\tau_m = 0$) at distance $d_{m,i} > d^S$ with no spillover effect:*

$$y_{mt}(0, \sigma_m(\tau)) = y_{mt}(0, \sigma_m(\mathbf{0})) \tag{29}$$

Using binary variable $s_m \in \{0,1\}$ to indicate whether untreated units ($\tau_m = 0$) are exposed ($s_m = 1$; $d^T < d_{m,i} \leq d^S$) or not ($s_m = 0$; $d_{m,i} > d^S$) to spillovers from treated units, OLS estimation of (11) delivers consistent estimates of the average treatment effect on the treated δ^T ($\tau_m = 1$) and of the average spillover effect on the close-to-treated (exposed to spillovers) δ^S ($s_m = 1$) under the following assumptions:

Assumption 3. *Parallel trends in treated ($\tau_m = 1$) and control ($s_m = 0$):*

$$E\left[y_{m1}(0, \sigma_m(\mathbf{0})) - y_{m0}(0, \sigma_m(\mathbf{0})) \,|\tau_m = 1\right] = E\left[y_{m1}(0, \sigma_m(\mathbf{0})) - y_{m0}(0, \sigma_m(\mathbf{0})) \,|s_m = 0\right] \tag{30}$$

Assumption 4. *Parallel trends in close-to-treated ($s_m = 1$) and control ($s_m = 0$):*

$$E\left[y_{m1}(0, \sigma_m(\mathbf{0})) - y_{m0}(0, \sigma_m(\mathbf{0})) \,|s_m = 1\right] = E\left[y_{m1}(0, \sigma_m(\mathbf{0})) - y_{m0}(0, \sigma_m(\mathbf{0})) \,|s_m = 0\right] \tag{31}$$

To test the parallel trends assumptions (30) and (31), we take changes in the outcome variables from the decade preceding the opening of the plants (2001-2011) and regress them on the vector of dummy variables equalling 1 if a waste management plant opens during the period 2016-2019. We perform the parallel trends test on the 'central' rings specification from table 2.[13] The results in table A.1 show that the decade-long changes in the outcome variables *before* treatment occurs are not significantly different between the two inner rings (0-4 km and 4-8 km) and the outer control ring (8-30 km).

Table A1: Pre-treatment trends (2001-2011$^{\diamond}$ change)

	(1)	(2)	(3)	(4)	(5)
	0-4km	4-8km	8-30km	(1)-(3)	(2)-(3)
Δshare(left vote)	0.031	0.025	0.025	0.006*	0.000
	(0.046)	(0.034)	(0.041)	(0.003)	(0.002)
Δln(domestic property price)	0.190	0.154	0.157	0.033	-0.003
	(0.236)	(0.245)	(0.242)	(0.019)	(0.019)
Δln(business property price)	0.061	0.045	0.047	0.014	-0.002
	(0.231)	(0.208)	(0.202)	(0.021)	(0.016)
Δln(income p.c.)	0.050	0.044	0.049	0.001	-0.005
	(0.060)	(0.058)	(0.064)	(0.005)	(0.004)
Δln(taxpayers > €55k)	0.673	0.679	0.666	0.007	0.013
	(0.280)	(0.309)	(0.314)	(0.021)	(0.021)

Notes: Standard errors clustered by plant site. *** : p value < 0.01; ** : p value < 0.05; * : p value < 0.10. $^{\diamond}$: property prices: 2004-2011; left vote share: 2001-2008 (parliamentary election results at the municipal level).

[13] Similar results arise when the other rings specifications are employed.