INEFFICIENCIES IN LOCAL INFRASTRUCTURE: EVIDENCE FROM DRINKING WATER IN CALIFORNIA

PIERRE BODÉRÉ, GIULIA BRANCACCIO AND KARAM KANG

Abstract. Local elected officials oversee critical infrastructure investments, yet limited electoral competition may weaken the alignment between local policy and constituent welfare. Using new data on system performance and investment for California's drinking-water systems, we document that weaker electoral competition is associated with higher water pollution and a muted response to public investment aid. We then build and estimate a dynamic model in which residents differentially value water quality and vote accordingly, while office-seeking politicians weigh private costs against re-election prospects. Residents' preferences are identified from housing transactions at water-system boundaries and electoral incentives are identified by exploiting variation in voter composition. Counterfactual increases in accountability reallocate investment toward high-need systems, reduce pollution, and generate progressive welfare gains. Limited accountability also dampens the effectiveness of federal and state subsidies, diverting funds from high-return investments and reducing the overall impact of infrastructure policy.

1. Introduction

Infrastructure investment in the United States has become a growing concern as spending struggles to keep pace with an aging system. Local governments are responsible for roughly half a trillion dollars annually in transportation and water infrastructure, yet accountability of their elected officials is often questioned (Berry, 2009). When electoral oversight is limited, officeholders may pursue objectives that diverge from constituent welfare, weakening the link between local needs and the allocation of public investment. This paper quantifies how electoral accountability

Date: March 1, 2026.
We thank audiences at NYU, UVA, UBC, UIUC, Tufts, Wisconsin-Madison, UCL, Princeton, IIOC, Duke, Stanford, Berkeley Haas, SED, WIEM, Chicago IO Day, UEA, Yale SOM, HBS, Penn State, WUSTL for helpful comments and suggestions. We are grateful to Arie Beresteanu, Ashley Langer, and Lindsey Currier for their insightful discussions. We thank Alejandro Herrera Caceido, Jiaoyang Li, and Genna Liu for excellent research assistance.
Bodéré (email: pierre.bodere@yale.edu): Yale School of Management.
Brancaccio (email: giulia.brancaccio@nyu.edu): New York University, CEPR, and NBER.
Kang (email: karam.kang@wisc.edu): University of Wisconsin-Madison and NBER.

affects the efficiency of infrastructure allocation and the effectiveness of large-scale public subsidies, and evaluates the resulting welfare implications.

We study drinking water infrastructure in California. Despite substantial federal and state funding, including $41.1 billion allocated since 1996 through Safe Drinking Water Act programs, chronic underinvestment persists, with measurable health consequences (Allaire et al., 2018; DiSalvo and Hill, 2024; Keiser et al., 2023). These challenges suggest that the problem goes beyond fiscal capacity. Moreover, most drinking water systems are governed by locally elected officials, and they tend to face little electoral competition: over the period of our study (2003-2022), 23% of city-council, 15% of county-supervisor, and 66% of water-district board races lacked challengers. Even when contested, incumbents rarely face narrow margins.

We assemble a unique, comprehensive dataset linking drinking water system investments, contaminant readings and violations, investment decisions and costs, election outcomes, and housing transactions for California jurisdictions. We first document that jurisdictions with stronger electoral competition—measured by contested-election frequency, narrow vote margins, or election timing coinciding with federal elections ("on-cycle" elections, as opposed to "off-cycle" ones)—exhibit significantly lower pollution levels, even after controlling for demographics, observed water system needs, and county and year fixed effects. Exploiting staggered transitions from off-cycle to on-cycle elections, we also find sustained reductions in contamination. In addition, using time-series variation in federal and state funding expansions and an exposure measure of expected intensity of receiving such funding, we show that exogenous funding opportunities increase investment probability, and that this response is stronger where elections are more competitive, indicating that accountability improves the alignment between external funding and local welfare gains.

In light of these findings, we develop and estimate a dynamic model of office-seeking policymakers who choose infrastructure investment each period. Investment improves water quality with persistent but depreciating effects and must be financed through subsequent water-rate increases. Residents, who differ in their willingness to pay for quality improvements, cast ballots based on period-to-period changes in their own utility, provided that the election is contested and they turn out. The long-lived policymaker thus faces a dynamic trade-off: investing can raise re-election prospects by improving constituents' welfare in the future, but may entail substantial private costs, potentially diverting the policymaker's attention from other policy-relevant issues. Per-period payoffs balance residents' preferences against these costs, with the

weight placed on residents' welfare determined by the electoral incentives and the institutional features that shape them—such as the likelihood of contested elections, relative turnout rates, and the salience of water issues.

Identification requires separating policymakers' objectives from voters' preferences and, within policymakers' payoffs, disentangling electoral incentives from private costs of undertaking investment. We estimate heterogeneous residents' willingness to pay for water quality by combining a housing choice model with a boundary discontinuity design that leverages discontinuities at water system boundaries, following Bayer, Ferreira and McMillan (2007). We measure the causal impact of investment on pollution and water rate using an event-study design. Together, these components identify the voters' welfare gains of infrastructure investment.

To isolate electoral incentives within policymakers' objectives, we exploit variation in the responsiveness of investment to voter composition. Voter preferences for water quality depend on demographic characteristics, and electorate composition varies both within and across jurisdictions. Within jurisdictions, shifts from off-cycle to on-cycle elections alter turnout and thus the identity of the decisive electorate. Using individual-level turnout data, we trace how these sources of variation affect investment decisions, thereby identifying the weight policymakers place on voter welfare separately from their own private costs or preferences.

Our structural estimates indicate that investments reduce water quality violations by 27% in years three through eight following investment. On average, investments increase annual household water rates by approximately $24 (2020 dollars) over the first decade, with substantial heterogeneity across systems depending on their typical financing structure for investment. Residents' median willingness to pay for a one-standard-deviation improvement in water quality is $66 per household per year. Combining estimated benefits and costs reveals substantial underinvestment: for roughly 39% of governments—serving 64% of Californians—additional investment yields positive net benefits averaging $42 per household annually, with benefits approximately 4.5 times costs. We find that this underinvestment is more prevalent where electoral competition is weaker.

Our results show that politicians do respond to resident preferences, but their investment decisions are constrained by limited administrative and technical capacity. The latter constraints are higher in smaller water systems and in governments with fewer highly paid, skilled employees capable of managing complex capital investments.

At the same time, the weight politicians place on residents' welfare is significantly higher in governments that hold contested elections more frequently.

We use our estimated model to quantify how local electoral accountability shapes infrastructure investment, welfare, and the effectiveness of federal and state subsidies. Accountability enters the model through the weight placed on residents' welfare. We consider a counterfactual in which accountability is raised to a plausible upper bound, corresponding to governments that faced consistently contested elections, and solve for the resulting investment policies. The counterfactual reallocates investment toward systems with the greatest underlying pollution, aligning capital spending more closely with locally specific marginal welfare gains. In the most pollution-prone quintile, investment probabilities rise by more than 25 percent relative to baseline levels, leading to pollution reductions of up to 0.35 standard deviations. By contrast, investment declines modestly in cleaner systems where marginal benefits are lower. These changes translate into economically meaningful welfare gains: in the most pollution-prone systems, average household welfare increases by up to $20 per year.

Greater accountability also strengthens the effectiveness of existing federal and state subsidies. In the counterfactual, pollution reductions from current subsidy programs increase by approximately 37 percent on average and by nearly 50 percent in the most polluted systems. Stronger electoral competition aligns subsidy take-up more closely with social returns. While decentralization is often justified by local governments' superior information about heterogeneous community needs (Oates, 1972), our results show that this informational advantage translates into efficient targeting only when policymakers are electorally disciplined. Electoral accountability therefore affects not only local investment choices but also enhances the return to large-scale federal and state water infrastructure programs, with larger gains in systems facing the highest environmental need.

1.1. **Literature review.** A growing literature studies whether decentralized infrastructure decisions are efficient, in the context of schools (Cellini et al., 2010; Biasi et al., 2025), ports (Brancaccio et al., 2024), high-speed rail (Fajgelbaum et al., 2023), drinking water systems (Sileo, 2023), and wind farm siting (Kashner, 2026). Measuring efficiency requires quantifying both the benefits and the costs of investment. We estimate benefits using housing-based measures of residents' willingness to pay for water quality, similar to Cellini et al. (2010), Biasi et al. (2025), and Kashner (2026).

On the cost side, we leverage detailed data on drinking water quality, water charges, and local governments' financing of infrastructure projects.

We move beyond measuring inefficiency to investigating its causes, with a focus on electoral accountability. We show that local electoral competitiveness affects both water quality and the effectiveness of federal subsidies, which motivates an examination of how *office-seeking* officials choose investments.[1] Whereas existing research highlights external constraints on public utilities—such as rate-of-return regulation (Lim and Yurukoglu, 2018; Gowrisankaran et al., 2024) or regulated pricing (Timmins, 2002)—we focus on distortions arising from politicians' welfare weights and capacity constraints in executing large capital projects.

A central contribution of the paper is the estimation of policymakers' objective functions from observed investment decisions. Our approach exploits variation in the welfare gains from investment across constituents, combining residents' willingness to pay with the effects of investment on water quality and rates. Cross-sectional variation arises from heterogeneous willingness to pay across income groups and differences in jurisdictions' income composition, in the spirit of Mian et al. (2010). Within jurisdictions, migration, electoral rule changes, and exogenous shocks to pollution shift the composition of participating voters and thus the aggregate welfare return to investment. Because investment decisions recur over time, this panel structure provides an additional source of variation to identify policymakers' responsiveness to constituent welfare. This approach complement strategies that rely on exogenous shifts in electoral incentives, such as term limits (Besley and Case, 1995; List and Sturm, 2006; Sieg and Yoon, 2017; Aruoba et al., 2019; Sieg and Yoon, 2022), anti-corruption reforms (Avis et al., 2018), or campaign dynamics (Iaryczower et al., 2024).

Closely related is Fajgelbaum et al. (2023), who study how political preferences shape the spatial design of California's high-speed rail. Using voting data and a quantitative spatial model, they recover political weights from observed policy choices. We share their revealed-preference approach to identifying political objectives. However, our setting differs along several dimensions. First, whereas they analyze a one-time statewide referendum and centralized project design, we study repeated decentralized

[1]Previous studies have focused on politician rent-seeking and favoritism (Avis et al., 2018; Burgess et al., 2015), myopic voters (Healy and Malhotra, 2009), information disclosure (Bennear and Olmstead, 2008), time consistency (Lim and Yurukoglu, 2018; Ouazad and Kahn, 2025), cross-jurisdiction spillover (Bordeu, 2023), constituent political preferences (Fajgelbaum et al., 2023), and debt market frictions (Adelino et al., 2017; Agrawal and Kim, 2022).

investment decisions by locally elected officials. Second, we directly quantify electoral accountability by estimating how policymakers' responsiveness to constituent welfare varies with electoral competition. Third, we show that imperfect accountability distorts not only investment allocation but also the effectiveness of public subsidies, expanding on recent research demonstrating the effectiveness of federally funded drinking water investments (Keiser et al., 2023). These distinctions allow us to analyze dynamic political agency and fiscal misallocation in routine public investment rather than a single high-salience project.

2. Drinking water infrastructure

Water utilities, typically run by local governments like cities, manage three critical operations. First, they source raw water by pumping it from surface water bodies like lakes and rivers, or from underground aquifers. Second, they process this water through treatment facilities to ensure it meets safety standards for human consumption. Third, they maintain storage systems and distribution networks to provide consumers with reliable, uninterrupted access to water. These operations depend on extensive infrastructure that demands ongoing financial investment and regular maintenance to function effectively.

To fund infrastructure investments, local governments can choose between debt financing or pay-as-you-go funding, with both methods primarily relying on revenue from water services. This stems from the structure of municipal water operations as independent enterprises, separate from general government operations. Unlike other government services supported by tax revenue, water services are funded through user charges, and hence any financing arrangements for water infrastructure projects are ultimately secured by these water-related revenues (Posenau, 2021). As a result, new infrastructure investments typically necessitate higher water charges.

Water charges comprise two parts: a fixed charge (including local taxes and fees) and a consumption-based service charge. In California, these charges face strict regulatory oversight. Proposition 218 empowers property owners to block water rate increases, while cost-of-service requirements limit price differentiation across households. These constraints result in simple pricing structures among water systems in our sample: 45% use uniform pricing (flat fixed and service charges), while 55% employ tiered systems with modest usage-based price variation. Even when water charges must increase to cover investments, regulatory requirements mandate that

revenue increases be distributed proportionally across rate components based on customer service levels.[2]

California's water services are managed by two types of local entities: cities and special districts, each governed by elected officials – mayors and council members for cities, and board members for water districts. The local nature of these races, however, raises issues of a lack of robust electoral accountability. Indeed, these elections consistently see voter participation rates at only half the level of national elections, with even lower turnout when they are not synchronized with presidential or gubernatorial races (Hajnal and Lewis, 2003). This pattern of low voter engagement has raised concerns that special interests or only a subset of the electorate wield disproportionate power in these elections and government decision-making (Anzia, 2011).

The 1974 Safe Drinking Water Act (as amended in 1996) regulates about 90 pollutants by maximum contaminant levels (MCL), and mandates timely public notifications of violations: within 24 hours for acute health risks like E.coli (Tier I), and within 30 days for all other MCL violations (Tier II). Additionally, the 1996 Amendments created the Drinking Water State Revolving Loan Fund, which offers subsidized loans for addressing critical health-related infrastructure needs. Communities that are designed as "disadvantaged" may qualify for additional financial support, including zero-interest loans or principal forgiveness, and in California, the designation is based on the median household income in a water system's service area.

3. Data

We address the lack of comprehensive drinking water infrastructure investment data by building an original database spanning 2003-2022. We enrich this dataset by integrating information on water quality metrics, utility water charges, municipal finances, local political conditions, housing market transactions, local demographic and socioeconomic characteristics, and climate to enable detailed analysis. Below, we describe our data sources and key variables.

3.1. **Drinking water infrastructure investment data.** Our novel dataset tracks California local governments' drinking water infrastructure investments, including investment frequency, funding sources, and user financial impacts. To our knowledge,

[2]This is referred to as the proportionality requirement of property related fees, stipulated in Proposition 218 (see Article XIII D Section 6 of the California Constitution). See, as an example, a recent water rate study for the city of Madera (Madera, 2022).

this provides the first comprehensive view of local government infrastructure investment behavior on drinking water, expanding on prior work by Posenau (2021) and Sileo (2023).[3]

The data construction draws primarily from California's Financial Transactions Reports (FTR), mandatory annual GAAP financial statements filed by local governments. These reports detail revenues and expenses, and we extract data specific to water enterprise operations. We augment this core dataset with two additional sources: California Debt and Investment Advisory Commission (CDIAC) records and EPA State Revolving Fund records acquired through FOIA requests. Overall, our data covers local governments that reported water enterprise revenues and operate public water systems during the study period, encompassing 273 cities (serving 20 million residents) and 551 special districts (serving 14 million residents).[4]

To identify infrastructure investment, we cross-reference indebtedness data from FTR with individual loan record from CDIAC to pin down debt financing. Moreover, we rely FTR income statements and DWSRF records to isolate external funding sources (i.e., federal and state government grants).[5] We identify an investment event when the combined total of debt financing and external funding per year exceeds a minimum threshold.[6] This approach relies on the premise that municipalities rarely fund infrastructure projects entirely from internal resources.[7]

Table 1 summarizes key variables for the cities and special districts separately. Based on our investment measure, Panel A reveals that infrastructure projects are largely financed by issuing municipal bonds, and that federal and state (subsidized) loans and grants, including those from EPA's Drinking Water State Revolving Fund,

[3]Posenau (2021) examines water utilities' debt contracts using our data sources, but doesn't investigate infrastructure investment choices. Sileo (2023) leverages Kentucky's Water Resource Information System (WRIS) to compile data on water systems and infrastructure projects, but does not study water service charges or the source of funding.

[4]There are 6 county governments that satisfy the criteria, and we drop them in our analysis.

[5]Specifically, FTR and CDIAC detail each debt's principal amount, issuance year, purpose, designated repayment revenue streams, and any federal/state government involvement. For external funding, we combine intergovernmental transfer records from FTR water enterprise income statements with EPA's DWSRF loan and subsidy records.

[6]We set thresholds at $54,000 for cities and $35,000 for special districts (in 2020 CPI-adjusted dollars) to ensure we capture substantive municipal projects. Our main findings remain robust across various threshold specifications.

[7]Our analysis of post-2017 FTR data for cities, which records capital investment through equity for water enterprises, confirms our premise: the probability of an above-mean-sized investment occurring without debt financing or external funding is just 1.3%.

TABLE 1. Summary statistics

	Cities	Special districts
A. Infrastructure investment		
Annual investment frequency	0.14	0.12
Conditional on any investment		
Amount debt-financed or externally funded		
Mean (in $ millions)	43.7	16.7
Median (in $ millions)	6.7	4.1
Per household, Mean (in $)	1,278	2,438
Per household, Median (in $)	466	473
Financing or external funding sources		
% financed via municipal bonds	70.15	61.22
% financed via federal/state loans	19.76	11.16
% externally funded (i.e., grants)	10.09	27.62
B. Annual cost of water per household		
Mean (in $)	753 (673)	909 (929)
Median (in $)	546	566
C. Water pollution		
Pollution measure > 0	0.260	0.199
EPA - Tier I	0.005	0.008
EPA - Tier II	0.065	0.084
Number of governments	273	551

Notes. This table presents summary statics of key variables used in our analysis, based on the 273 cities and 551 special districts that provide drinking water services, for the period of 2003-2022. The mean values are presented with standard deviations in parentheses, unless stated otherwise. All investment amounts and rates are adjusted to 2020 dollars using CPI. †: These conditional statistics are for municipalities that made at least one investment during the study period, and are based on the *total* debt financing and external funding across all investments per municipality.

remain a secondary source of funding.[8] This comprehensive data on investment, therefore, allows us to complement the existing literature that looks at the role of EPA's Drinking Water State Revolving Fund in improving access to clean water.

Moreover, Panel A in Table 1 shows that investment is relatively frequent, with 14% of cities and 12% of special districts investing in any given year. Among investing municipalities, the median funding secured through debt financing and external sources reached $6.7 million for cities and $4.1 million for special districts. However, mean values demonstrate substantial right-skew, at $43.7 million and $16.7 million

[8]Municipalities often fund infrastructure projects through a mix of internal funds, debt-financing and external funds like grants. We find that debt instruments like municipal bonds and government loans comprise 67.3% of the infrastructure investment cost, based on post-2017 FTR city data. This aligns with practitioner estimates of 65% debt financing (Hansen and Mullin, 2022).

respectively. This disparity persists when examining per-household investment: in special districts, the median investment of $466 per household represents much less than the mean value, $1,278.

As described above, local governments finance water infrastructure investment through user fees and charges. Therefore, to assess the financial burden on water service users, we calculate the per-household cost by dividing the annual total of sales, taxes, and fees (from FTR water enterprise income statements) by the number of households in each municipality's water service area.[9] Panel B in Table 1 reveals a median annual cost of $546 per household for cities and $566 for special districts, with means exceeding these values in both cases. Notably, these costs show substantial variation across municipalities, as evidenced by large standard deviations.

3.2. **Additional data sources.** We measure the level of drinking water quality and service at the public water system level. The CA Division of Drinking Water provides pollution measurements for more than 70 federally regulated contaminants. In addition, the EPA website supplies records of Safe Drinking Water Act violations. We link water systems to their municipalities through manual name verification, online research, and a database on California community water system institutional types (Dobbin et al., 2023).

Panel C in Table 1 displays summary statistics for drinking water pollution in California systems. EPA violations happen on a regular basis: 6.5% of cities and 8.4% of special districts face Tier II EPA violations in any given year. However, more severe (Tier I) violations are quite rare, with a frequency below 1% for both cities and special districts.

Given that water pollution is measured in multiple dimensions, we aggregate these measurements into an index. The water pollution at system i during year t, denoted by q_{it}, is the sum of an indicator for EPA violations and an indicator for each pollutant c's concentration reading (q_{cit}) above the maximum contaminant level (MCL_{ct}), divided by the total number of pollutants plus one:

$$q_{it} = \frac{\mathbf{1}\{\text{EPA violation}_{it}\} + \sum_c \mathbf{1}\{q_{cit} > MCL_{ct}\}}{1 + \#\text{ pollutants measured}_t}. \tag{1}$$

[9]We are currently collecting more precise measures of resident water fees and charges by directly contacting individual water utilities to obtain current rate schedules and documentation from rate hearing proceedings.

Here we exclude disinfection byproducts, and thus the denominator for this measure is around 71. Unless noted otherwise, we use this measure throughout the study.[10] We find that in any given year, 26% of cities and 20% of special districts reach measures of pollution in excess of regulated MCL.

To study the incentives of local government to invest in drinking water infrastructure, we combine our investment data with electoral data from several sources. City election information—including dates, turnout, and competitiveness—comes from the California Election Data Archive (CEDA), while special district election data is being manually collected from county election websites. To analyze mayoral term limits, we merge election results with the International City/County Management Association's (ICMA) Municipal Form of Government Surveys, which document both government structure (mayor-council or council-manager) and term limit policies.[11]

To analyze spatial patterns, we integrate multiple datasets with water service area boundaries from the California State Water Resources Control Board: demographic and socioeconomic characteristics from the American Community Survey and climate data. This spatial mapping enables analysis at both the water system and municipality levels. Finally, to understand residents' preferences regarding water quality, we analyze CoreLogic's housing transaction data, enriched with buyer demographics through linkage to Home Mortgage Disclosure Act (HMDA) records and tract-to-tract commuting patterns from Longitudinal Employer-Household Dynamics (LEHD) data.

4. Suggestive evidence for infrastructure under-investment

Using a novel, comprehensive dataset, this section presents two main results. First, we show that local water pollution is systematically related to features of the political process that shape electoral incentives but are plausibly orthogonal to the underlying costs and benefits of infrastructure investment. We find that weaker electoral competition is associated with lower drinking water quality, a result that remains robust when exploiting within-jurisdiction panel variation in election timing. Second, leveraging variation in the availability of federal and state funding, we present

[10]When the unit of analysis is the local government, we further aggregate this index for those that own multiple water systems, with the weight being the number of connections of each water system. In our sample, 89% of the local governments providing drinking water services to residents own a single water system.

[11]The next version of the paper will to incorporate L2 data to examine the demographic patterns in local turnout.

suggestive evidence that electoral conditions also attenuate the effectiveness of subsidies for infrastructure investment. Jurisdictions with weaker electoral competition respond substantially less to available funding, holding constant their underlying system needs. Taken together, these patterns point to important misallocation in the provision of safe drinking water, and suggest that electoral competition acts as a disciplining force that aligns official actions with their constituents' needs.

4.1. **Electoral competition and water pollution.** Electoral competition for seats on city councils or special-district water boards varies widely. In Santa Fe Springs, every city-council election from 2003–2022 was contested, and the average winning margin was 1.8%. By contrast, in the Ponderosa Community Services District, there were only two board elections over the same period, and the average winning margin exceeded 45%.[12] Such differences in electoral incentives may shape officials' investment decisions even when constituents' preferences are similar, generating misallocation and inefficiency in the provision of safe drinking water. Here we study whether this heterogeneity in competition accounts for the cross-jurisdiction differences in water quality documented in Section 3, while plausibly holding fixed investment costs and constituent preferences.

4.1.1. *Cross-sectional evidence.* We examine how drinking water pollution in jurisdiction i during year t, as defined in (1), varies with the jurisdiction's overall level of electoral competition, e_s. The specification includes jurisdiction and period fixed effects, μ_s and ϕ_t, respectively, as well as time-varying jurisdiction-level demographic characteristics, water system attributes, and precipitation, collected in the vector $\mathbf{x}_{s,t}$.

$$q_{s,t} \;=\; \beta_e\, e_s \;+\; \beta_x \mathbf{x}_{s,t} \;+\; \mu_s \;+\; \phi_t \;+\; \varepsilon_{s,t}, \tag{2}$$

We construct multiple measures of electoral competition. First, we measure election frequency by dividing the sample into two-year periods and recording whether an election occurred in each period. Our measure equals the fraction of two-year periods in which an election took place, capturing the regularity with which incumbents face electoral review. Second, we compute the average vote margin for incumbent candidates. This measure is available for jurisdictions whose election results are recorded in the CEDA database, typically cities, counties, and selected special districts. Third, similarly to the first measure, we calculate the fraction of two-year periods in which

[12]In many small districts, board seats go uncontested, and incumbents retain office by default. For example, a water district near Sacramento held its first election in 43 years in 2019.

TABLE 2. Electoral Competition matters for water quality

	Average % above MCL		
	(1)	(2)	(3)
Frequency of contested elections	-0.012*		
	(0.007)		
Average vote margins for incumbents $\leq 10\%$		-0.011**	
		(0.005)	
Frequency of on-cycle elections			-0.016**
			(0.008)
Median household income (standardized)	-0.009***	-0.015***	-0.009***
	(0.003)	(0.005)	(0.003)
County FE, Year FE	Yes	Yes	Yes
Number of observations	14,017	8,314	14,017
Mean of dependent var.	0.030	0.029	0.030
R^2	0.075	0.082	0.076

Notes: This table reports regression estimates where the dependent variable is pollution readings from public water systems, aggregated to the government-year level over the period 2003-2022. Standard errors, clustered at the government level, are shown in parentheses. Statistical significance is denoted by $^{*}p < 0.10$, $^{**}p < 0.05$, and $^{***}p < 0.01$. †. Additional controls include population served, primary water source (surface vs. groundwater), socio-economic covariates (% white, population density), the fraction of readings conducted in each calendar month, and precipitation measures (mean, 75^{th} percentile, 90^{th} percentile, and the fraction of days with no rain).

an on-cycle election was held, where "on-cycle" elections are those held on federal Election Days.

Table 2 shows that electoral competition is negatively associated with drinking water pollution. Across specifications, measures of electoral competition—whether captured by the frequency of contested elections, narrow incumbent vote margins, or the frequency of on-cycle elections—enter with negative and statistically significant coefficients, even after controlling for variables that affect the underlying costs and benefits of infrastructure investment and water service. The magnitudes are economically meaningful. The implied reduction in pollution from increasing election frequency from 0 to 1 is comparable to the reduction associated with a one standard deviation increase in median household income. In other words, moving from no elections to elections in every two-year period yields an improvement in drinking water quality similar in size to that generated by a substantial increase in local income, highlighting the quantitative importance of electoral incentives. These results suggest that electoral competition acts as a disciplining force that aligns official actions with constituent needs. By narrowing the gap between official actions and

constituent preferences, high-frequency and high-stakes elections appear to mitigate misallocation in the provision of safe drinking water.

4.1.2. *Event study evidence.* A causal interpretation of the patterns in Table 2 is limited because electoral competition is an equilibrium outcome. For example, poor services can trigger backlash and attract more candidates, increasing competition. To address this concern, we exploit within-jurisdiction changes in local election timing. The literature shows that moving local elections on-cycle approximately doubles turnout and yields electorates that are more representative by race, age, and partisanship (Hajnal and Lewis, 2003; Hajnal et al., 2022).[13] Moreover, our data shows that on-cycle elections are more more likely contested and also have a lower vote margin.[14]

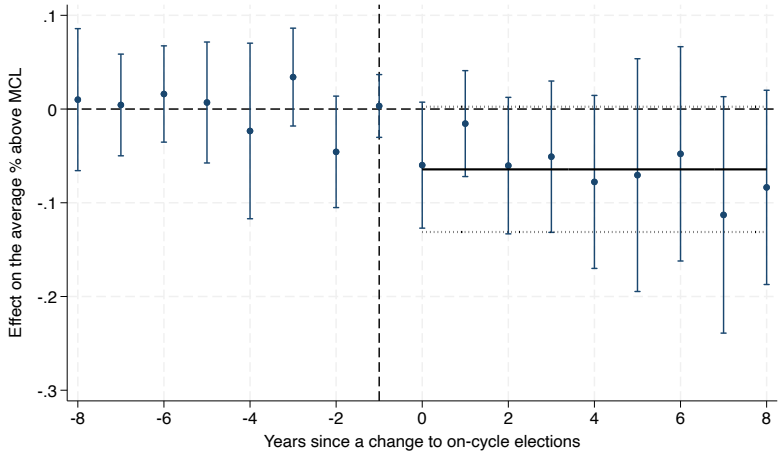
Many jurisdictions have shifted their election dates to coincide with federal contests—especially following the California Voter Participation Act, which required some local governments to move on-cycle from 2018. In our sample, 58 governments switched their election schedule from off-cycle to on-cycle prior to 2018, and an additional 57 governments switched in or after 2018, while 85 governments in our sample continued to hold elections off-cycle. Leveraging these jurisdictions and the staggered adoption of on-cycle elections, we study the effects of election timing on local drinking water quality. To accommodate potentially heterogeneous treatment effects, we employ the doubly robust difference-in-differences methods developed by Sant'Anna and Zhao (2020) and Callaway and Sant'Anna (2021).

Figure 1 presents estimates from the event-study analysis, where the dependent variable is local drinking water pollution (q_{it} in (1)) and the policy event is the local legal change in election timing from off-cycle to on-cycle. The results indicate improvements in drinking water quality following the transition to on-cycle elections. The average treatment effect across the eight post-treatment years is -0.064 (government-level clustered standard error 0.034) units, equivalent to roughly half of the outcome's standard deviation. This effect is economically meaningful and statistically significant at conventional levels (p-value $= 0.059$), with a 95% confidence interval that narrowly includes zero. Importantly, there is no evidence of differential pre-trends prior to the

[13]Our data confirm this result. For example, council-member elections exhibit 27% turnout when aligned with federal or state elections, compared to 16% when held off-cycle.

[14]These findings contrast with de Benedictis-Kessner (2018), who, studying mayoral contests, shows that on-cycle elections confer a substantially larger incumbency advantage than off-cycle elections. A key distinction is that our balanced panel includes years with uncontested races, aligning with our focus on measuring electoral competition by whether races are contested and average winning margins.

FIGURE 1. Election Timing and Drinking Water Pollution



Notes: This figure presents estimates from the event-study specification for local drinking water pollution, as defined in (1), using the methods developed by Sant'Anna and Zhao (2020) and Callaway and Sant'Anna (2021). The policy event is the local legal change in election timing from off-cycle to on-cycle. Dots denote point estimates, and vertical lines indicate 95% confidence intervals based on standard errors clustered at the government level. Coefficients are reported for eight years before and eight years after the policy change. The solid horizontal line represents the average treatment effect across all eight post-treatment years, while the dotted horizontal lines depict its 95% confidence interval.

shift to on-cycle elections: the average pre-treatment effect is close to zero, supporting the validity of the research design. Overall, moving elections on-cycle is associated with sustained reductions in drinking water pollution.

4.2. **Electoral competition and investment.** The results in Table 2 and Figure 1 show that greater electoral competition is associated with improvements in drinking-water quality. A natural question is whether this relationship operates through elected officials' investment behavior. A challenge here is that assessing how electoral competition influences infrastructure investment requires quantifying both its benefits and its costs, as greater spending is not necessarily welfare improving. As preliminary evidence, we examine how officials respond to variation in federal and state funding availability, which lowers the effective cost of investment, and whether this response varies with the degree of electoral competition.

Figure A2 in the Appendix illustrates substantial temporal variation in realized grant funding, corresponding closely to major statutory funding expansions. With that, we exploit time-series variation in aggregate funding, forming funding shocks to comparable governments and combining these shocks with a predicted exposure

TABLE 3. External Funding Shocks and Infrastructure Investment

	Any investment		
	All	Weakly contested	Highly contested
	(1)	(2)	(3)
Expected grant intensity	0.3903***	0.3298*	0.4824**
	(0.1421)	(0.1782)	(0.2311)
Government FE, Year FE	Yes	Yes	Yes
Observations	14,320	7,160	7,160
R^2	0.0602	0.0616	0.0621

Notes: The dependent variable is an indicator for whether government i undertakes a drinking water infrastructure investment in year t. Standard errors, clustered at the government level, are reported in parentheses. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$. The key independent variable is expected grant intensity, z_{it}, defined in the text as a shift-share measure combining cross-sectional exposure to funding with time-series shocks to comparable governments. All specifications include government and year fixed effects. Column (1) uses the full sample. Columns (2) and (3) split the sample at the median election frequency into weakly contested and highly contested governments, respectively.

measure to state and federal funding. Specifically, we define the expected grant intensity for government s in year t, z_{st}.[15] First, we estimate government-level exposure to federal and state funding, p_s, by predicting the probability of receiving external funding using a random forest based on observed attributes that affect funding priority, including water system size, median household income, and measures of fiscal health, among others. Second, we construct funding shocks, g_{st}, using realized per capita grants among demographically similar governments. Governments are grouped into donor pools via k-means clustering on baseline (2003) population, income, and housing density, with a median pool size of 33. The expected grant intensity is then defined as $z_{st} = p_s g_{st}$.

Table 3 shows that greater funding availability increases the probability of investment within government. Column (1) indicates that a one-unit increase in z_{it} raises the investment probability by 0.39, implying a substantial response to external funding. When splitting the sample at the median election frequency into weakly and highly contested governments, Columns (2) and (3) show that the effect is larger among highly contested governments.

Taken together, our findings show that electoral competition improves drinking water quality and amplifies governments' investment responses to external funding

[15]See the Appendix for a more detailed description of the construction of z_{st}.

shocks. However, because both benefits and costs of infrastructure investment are difficult to measure directly, the evidence in this section alone cannot determine whether, and how much, stronger electoral incentives enhance welfare or investment choices. We therefore develop and estimate a structural model to quantify the net benefits of infrastructure investment and to assess how electoral incentives affect the efficiency and to explore policy implications for infrastructure provision.

5. Model

This section develops a model of an elected governing body's dynamic investment decisions in managing a jurisdiction's public drinking-water system. The body—e.g., a city council or special district board—comprises several members who face reelection on a staggered, two-year cycle, so we index periods by t in two-year increments. The jurisdiction's residents differ in how they value the trade-off between water quality and rates, and they decide whether to retain or replace the members standing for election based on current and past outcomes in these dimensions. Because a subset of seats is contested in any period, we model the governing body as a single decision maker solving an infinite-horizon problem, hereafter referred to as the *policymaker*.[16]

Each resident i is characterized by $\boldsymbol{y}_i$, which includes dwelling type (single- vs. multi-family, y_{1i}), income, and race. Let $F_{\boldsymbol{y},t}$ denote the cross-sectional distribution of $\boldsymbol{y}$ in the jurisdiction in period t.[17] Let $\boldsymbol{x}_t \in \mathcal{X}$ collect exogenous state variables that affect water quality and costs—e.g., system size and water source—as well as political features, such as jurisdiction type (city, dependent special district, independent special district) and election timing (on- vs. off-cycle).

At the start of each period t, the policymaker chooses whether to undertake a capital investment, setting $a_t = 1$ if they proceed and $a_t = 0$ otherwise. During the period, the realized water quality $q_t \in \mathbb{R}_+$ and the average per-resident cost of service $\bar{r}_t \in \mathbb{R}_+$ (covering capital, operations, and maintenance) are publicly observed. Given (i) a balanced-budget requirement that user revenues cover per-period costs, (ii) institutional and legal constraints (see Section 2), and (iii) inelastic demand, a

[16]The governing body is a committee whose members change over time, and we model that electoral risk and turnover scale continuation values rather than replacing the governing body entirely given that seats are staggered.

[17]We assume that this distribution evolves exogenously with respect to investment choices.

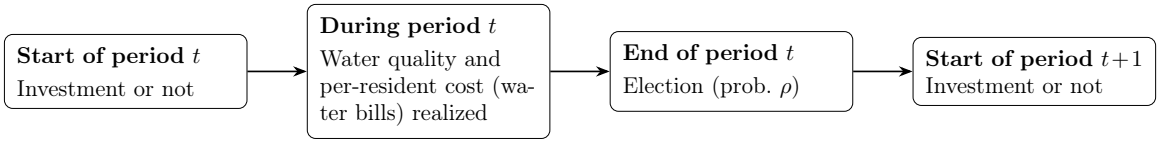


FIGURE 2. Timeline

known pricing rule below maps $(\bar{r}_t, y_{1i})$ into resident i's bill.[18] Specifically,

$$r_{it} \equiv \gamma(y_{1i}, \boldsymbol{x}_t)\bar{r}_t. \tag{3}$$

At the end of the period, an election occurs with probability ρ_t. If no election is held, the governing body's composition remains unchanged; otherwise, the next period's membership reflects the vote outcome. Figure 2 summarizes the timeline.

5.1. **Voting and election outcomes.** When an election occurs, resident i turns out with probability $\pi(\boldsymbol{y}_i, \boldsymbol{x}_t)$; otherwise, she abstains. Abstracting from seat-specific contests within period t, we treat the event as a single election that either approves or disapproves the policymaker.

We parameterize resident i's trade-off between water quality and rates by $\omega(\boldsymbol{y}_i) \in \mathbb{R}_+$, interpreted as her marginal willingness to pay for quality. Using the rate schedule (3), the per-period utility from water services is

$$u(q_t, \bar{r}_t, \boldsymbol{y}_i) \;=\; \omega(\boldsymbol{y}_i)\, q_t \;-\; \gamma(y_{1i})\bar{r}_t.$$

Conditional on turning out, j votes for the policymaker if her utility from current water services is at least her benchmark utility plus an idiosyncratic preference shock for the challenger (e.g., charisma or valence), v_{it}:

$$u(q_t, \bar{r}_t, \boldsymbol{y}_i) \;\geq\; v_{it}.$$

The shock v_{it} is drawn from $F_{v(\boldsymbol{y}_i)}$, independent of water investments and outcomes.

Let $\tilde{\pi}(\boldsymbol{y}, \boldsymbol{x}_t) \equiv \pi(\boldsymbol{y}, \boldsymbol{x}_t) \big/ \int \pi(\boldsymbol{y}', \boldsymbol{x}_t)\, dF_{\boldsymbol{y}t}(\boldsymbol{y}')$ denote turnout weights. The vote share for the incumbent is

$$\tilde{\mu}_t \;=\; \int \tilde{\pi}(\boldsymbol{y}, \boldsymbol{x}_t)\, F_{v(\boldsymbol{y})}\big[u(q_t, \bar{r}_t, q_{t-1}, \bar{r}_{t-1}, \boldsymbol{y})\big]\, dF_{\boldsymbol{y},t}(\boldsymbol{y}).$$

Aggregate uncertainty in translating vote share into victory is captured by $\eta_t \sim F_\eta$ with mean zero; the incumbent wins iff $\tilde{\mu}_t - \eta_t \geq \frac{1}{2}$, so the winning probability is

[18]We thus abstract from an explicit rate-setting problem for the policymaker. Typical water rate schedules depend on dwelling type and metered consumption, with the latter exhibiting limited variation within a given jurisdiction and dwelling type.

$F_\eta\big(\tilde{\mu}_t - \frac{1}{2}\big)$. If $\epsilon_{it} \sim \text{Unif}\Big[-\frac{1}{2\sigma_\epsilon}, \frac{1}{2\sigma_\epsilon}\Big]$ and $\eta_t \sim \text{Unif}\Big[-\frac{1}{2\sigma_\eta}, \frac{1}{2\sigma_\eta}\Big]$, then the winning probability is

$$\mu(q_t, \bar{r}_t, q_{t-1}, \bar{r}_{t-1}) \;=\; \frac{1}{2} \;+\; \sigma_\eta \int \tilde{\pi}(\boldsymbol{y}, \boldsymbol{x}_t)\, \sigma_\epsilon(\boldsymbol{y})\, u(q_t, \bar{r}_t, q_{t-1}, \bar{r}_{t-1}, \boldsymbol{y})\, dF_{\boldsymbol{y},t}(\boldsymbol{y}), \quad (4)$$

provided the arguments lie within the corresponding supports.

5.2. **Water quality and rate dynamics.** An important endogenous state is the number of periods since the last investment, or *system age* l_t. Given the investment choice a_t, it evolves deterministically as

$$l_t \;=\; \begin{cases} l_{t-1} + 1, & \text{if } a_t = 0, \\ 0, & \text{if } a_t = 1. \end{cases}$$

Intuitively, a younger system (smaller l_t) is more likely to deliver high-quality water.

We specify the law of motion for water quality as

$$q_{t+1} \;=\; \phi_q(\boldsymbol{x}_t, l_t, \zeta_{q,t})\,, \quad (5)$$

where the shock $\zeta_{q,t}$ drawn from F_{ζ_q} represents transitory influences on raw water quality and is assumed to be serially uncorrelated.

We model the per-period, per-resident cost of service as

$$\bar{r}_{t+1} \;=\; \phi_r(\bar{r}_t, \boldsymbol{x}_t, a_t, \zeta_{r,t})\,, \quad (6)$$

where the shock $\zeta_{r,t} \sim F_{\zeta_r}$ reflects transitory operations/maintenance costs, and is assumed to be serially uncorrelated.

5.3. **The policymaker's investment decision.** Because the policymaker chooses the investment at the start of period t prior to the realization of water service outcomes, the public state at that time is $\mathbf{s}_t \;\equiv\; (\, q_{t-1},\; \bar{r}_{t-1},\; l_{t-1},\; \boldsymbol{x}_t\,)$.

The private cost of investment borne by the policymaker is

$$\kappa(\mathbf{x}_t) + \epsilon_t,$$

where ϵ_t is i.i.d. Type-I extreme value with unit scale, and the payoff from not investing is normalized to zero. This term reflects the cost of time and efforts required to plan and implement the investment, capturing the notion, for example, that a time-constrained politcymaker may need to divert attention from other pressing issues, such as employment or security, in order to manage improvements in water quality.

In addition, if an election occurs and the policymaker is approved (i.e., the members up for reelection are retained), they obtain an extra benefit $\nu_t \geq 0$. We assume that both the election-occurrence probability and the victory benefit depend on exogenous state variables: $\rho_t = \rho(\boldsymbol{x}_t)$ and $\nu_t = \nu(\boldsymbol{x}_t)$.

Let β denote the discount factor. Let us define the policymaker's overall weight on residents' (long-run) utility from water services as:

$$\lambda(\boldsymbol{x}_t) \equiv \nu(\boldsymbol{x}_t)\, \rho(\boldsymbol{x}_t)\, \sigma_\eta\, \sigma_\epsilon. \tag{7}$$

This weight increases with the value of reelection (ν_t), the likelihood of an election (ρ_t), the sensitivity of victory to vote share (σ_η), and the salience of water service outcomes in voting (σ_ϵ). The expected value function is

$$\begin{aligned} V(\mathbf{s}_t) = \; \mathbb{E}_{\epsilon_t}\Bigg(\max_{a \in \{0,1\}} \Bigg\{ \lambda(\boldsymbol{x}_t) \int \tilde{\pi}(\boldsymbol{y}, \boldsymbol{x})\, \mathbb{E}[u(q_t, \bar{r}_t, q_{t-1}, \bar{r}_{t-1}, \boldsymbol{y}) \,|\, a, \mathbf{s}_t] \; dF_{\boldsymbol{y},t}(\boldsymbol{y}) \\ - \; a\,(\kappa(\mathbf{s}_t) + \epsilon_t) \; + \; \beta\, \mathbb{E}[V(\mathbf{s}_{t+1}) \,|\, a, \mathbf{s}_t] \Bigg\} \Bigg). \end{aligned} \tag{8}$$

Here, $\mathbb{E}_{\epsilon_t}(\cdot)$ averages over the contemporaneous shock ϵ_t; the inner expectations integrate over $(q_t, \bar{r}_t)$ induced by $(a, \mathbf{s}_t)$ and over the next-period state $\mathbf{s}_{t+1}$, respectively.

6. Identification and estimation procedure

Jurisdictions (cities and special districts) are indexed by s. In each period t, the policymaker in jurisdiction s chooses an investment $a_{st} \in \{0, 1\}$. Residents i have observed attributes $\boldsymbol{y}_i$; let $F_{\boldsymbol{y},s,t}$ denote the cross-sectional distribution of $\boldsymbol{y}$ within jurisdiction s at time t. Data provide, for each jurisdiction (city or special district) s and period t, the investment decision $a_{s,t}$ (from which the elapsed time since the last investment, $l_{s,t}$, is constructed), water quality $q_{s,t}$, the average per-resident cost $\bar{r}_{s,t}$, and jurisdictional covariates $\boldsymbol{x}_{s,t}$. In addition, we observe voter turnout, election occurrences, water rate schedules, aggregate water consumption by dwelling type, and government finances regarding loans and grants. Finally, we gather data on housing transactions and how the corresponding prices vary with water quality.

6.1. **Overview.**

We design our estimation approach to recover directly from the politician's behavior the weights they put on their constituents' utility, with minimal assumptions on the specific mechanism. To do that, we focus on the composite object $\lambda(\boldsymbol{x})$, and estimate it as a function of the election probability $\rho(\boldsymbol{x})$

The structural primitives fall into three groups. On the resident side, the primitives are the marginal willingness to pay for water quality, $\omega(\boldsymbol{y})$; the turnout, $\pi(\boldsymbol{y}, \boldsymbol{x})$; and the water pricing schedule $\gamma(y_1, \boldsymbol{x})$. On the government side, the primitives are the policymaker's private cost of investing, $\kappa(\boldsymbol{s})$, the weight they place on the constituents' utility $\lambda(\boldsymbol{x})$, and the intertemporal discount factor, β. Finally, the state transition primitives consist of $\psi_q(\boldsymbol{x}, l)$ and $\psi_r(\boldsymbol{x}, a)$, F_{ζ_q} and F_{ζ_r}, and the distribution of $\boldsymbol{x}_{s,t+1}$ conditional on $\boldsymbol{x}_{s,t}$.

Our estimation proceeds in three steps. Here we provide a brief overview and then subsequent sections describe each step of the estimation in detail. In the first step, we recover the election probability function $\rho(\boldsymbol{x})$ and the turnout function $\pi(\boldsymbol{y})$ from election occurrences and individual-level turnout records, and recover each jurisdiction's pricing rule $\gamma(y_1, \boldsymbol{x})$ using observed rate schedules, aggregate consumption by dwelling type, and average per-resident cost data. We also recover the state transitions for water quality and rate from the panel of $\{q_{s,t}, \bar{r}_{s,t}, a_{s,t}, l_{s,t}, \boldsymbol{x}_{s,t}\}$ as well as data on grants and loans.

Next, we recover the residents' marginal willingness to pay for water quality, $\omega(\boldsymbol{y})$. Note that jurisdiction-level investment decisions alone do not disentangle residents' preferences from policymakers' objectives, and election outcomes confound the marginal willingness to pay $\omega(\boldsymbol{y})$ with issue salience, given that micro-level voting data availability is limited for local elections. We therefore use auxiliary housing transactions data and identify $\omega(\boldsymbol{y})$ from household sorting: specifically, a boundary discontinuity design (Black, 1999) nested within an equilibrium housing-demand framework following Bayer et al. (2007).

In the last step, we recover the policymaker's private cost of investing, $\kappa(\boldsymbol{x})$, and the weights $\lambda(\rho(\boldsymbol{x}))$ from the observed investment decisions. We set $\beta = 0.92$, as the discount factor is not identified from choice-state data without additional restrictions (Rust, 1987; Magnac and Thesmar, 2002).

6.2. **Water quality transition estimation.** The key parameter in $\phi_q(\cdot)$ is the extent to which investment in water infrastructure can improve water quality over time. To partially address that $l_{s,t}$ is endogenous, we use the stacked event study approach to assess the impact of an investment across all investment events (Cengiz et al., 2019). For each investment event, h, we create an event-specific dataset, which includes observations from the treated government around the investment—between 3

years before and 16 years after the year of investment—and observations from all governments that did not have any major investments over the event window.[19] Since many readings are zero, we use the Poisson regression model with the conditional expectation as follows:

$$\mathbb{E}[q_{s,t,h}|\{I^{\tau}_{s,t,h}\}^{16}_{\tau=-3}, \boldsymbol{x}_{i,t}] = \exp\left(\sum_{-3}^{16} \phi^{q}_{\tau} I^{\tau}_{s,t,h} + \phi_x \boldsymbol{x}_{s,t} + \psi_{s,h} + \eta_{t,h}\right), \tag{9}$$

where $q_{s,t,h}$ is the water quality measured at the drinking water facility of government s at year t, included in the h^{th} investment event dataset. The treatment dummy $I^{\tau}_{s,t,h}$ equals 1 if the investment h was made τ years from year t by government s, and the specification includes event-specific government fixed effects and event-specific time effects, to ensure that identification comes from within an event.

Having estimated the dynamic treatment effect of investment on pollution, we take two additional steps to discipline the law of motion of q. First, we estimate a panel model of expected pollution at the system-year level as a function of system fixed effects, year fixed effects, and time-varying characteristics such as weather conditions, abstracting from investment decisions. Second, we use the cumulative treatment effects implied by observed investments to back out the counterfactual pollution path attributable solely to exogenous characteristics, that is, the level to which pollution would converge in the absence of investment, α_{st}:

$$\mathbb{E}[q_{st}] = \exp\left(\psi_x \boldsymbol{x}_{s,t} + \mu_s + \eta_t\right) = \exp\left(\alpha_{st} + \sum_{h} \sum_{\tau=0}^{t-t_h} \hat{\phi}^{q}_{\tau} I^{\tau}_{s,t,h}\right). \tag{10}$$

6.3. **Water cost transition estimation.** To estimate how investment affects per-resident water costs, which are then passed on to residents through water charges (Section 2), we combine a stacked event-study design, similar to that used in modeling water quality transitions, with additional structure to account for heterogeneity across governments and time. Specifically, for each investment event h we estimate

$$\log \bar{r}_{s,t,h} = \sum_{\tau=-3}^{16} \phi^{r}_{\tau} I^{\tau}_{s,t,h} + \mu_{s,h} + \rho_{t,h} + \gamma \boldsymbol{x}_{s,t} + \epsilon_{s,t,h}, \tag{11}$$

where $\bar{r}_{s,t,h}$ is the average water rate of government s in year t within the h^{th} event dataset, and $\gamma \boldsymbol{x}_{s,t}$ captures changes in ongoing maintenance and operating costs. To allow for heterogeneity in federal and state funding, which substantially reduces the

[19]Our results are robust to perturbations to the time window and weighting of the control group observations depending on the propensity score.

share of investment costs borne by residents, we estimate the water rate transition process separately for two groups defined by subsidy intensity (low and high). Subsidy intensity is predicted using the relationship between the observed subsidy share in realized investments and the government characteristics considered by federal and state authorities in allocating funds.

6.4. **Estimation of resident preferences for water quality.** To recover the residents' value for quality, we rely on house transaction prices in California from 2003 to 2019, and follow Bayer et al. (2007).[20] By nesting a boundary discontinuity design (Black, 1999) within an equilibrium housing demand model, we leverage variations in water quality across public water systems (PWS).[21]

Consider a potential buyer i with income y_{it} deciding between houses j in a local housing market in year t. Houses vary in prices p_{jt}, and in physical and surrounding neighborhood characteristics $\boldsymbol{x}_{jt}$. In addition, houses are potentially situated in different water systems, implying different water quality q_{jt} and rates r_{jt}. Potential buyers also consider the distance between each house j and their workplace, d_{ij}. Finally, houses differ according to an unobserved quality term ξ_{jt} potentially correlated with price. Together, these variables determine the utility U_{ijt} of buyer i for purchasing house j in year t:

$$U_{ijt} = \boldsymbol{\theta}_x \boldsymbol{x}_{jt} + \theta_q(y_{it}) q_{jt} - \theta_p(y_{it})(p_{jt} + r_{jt}) + \theta_d d_{ijt} + \psi_{b_j t} + \xi_{jt} + \varepsilon_{ijt} \tag{12}$$

where $\theta_q(y_{it}) = \theta_{q0} + \theta_{qy} \log(y_{it})$ and $\theta_p(y_{it}) = \theta_{p0} + \theta_{py} \log(y_{it})$ allow for heterogeneous price and pollution sensitivities for home buyers with different income levels. We assume that the house-specific taste shocks ε_{ijt} are distributed extreme value type I. With our notation, we can decompose this utility into three terms:

$$U_{ijt} = \delta_{jt} + \mu_{ijt}(\boldsymbol{\theta}_2) + \varepsilon_{ijt} \tag{13}$$

where δ_{jt} is the house-specific component how utility and $\mu_{ijt}(\boldsymbol{\theta}_2)$ contains the non-linear portion of utility (buyer–house specific). We denote the vector of linear parameters as $\boldsymbol{\theta_1} = (\boldsymbol{\theta_x}, \theta_{q0}, \theta_{p0})$.

[20]Relying on house prices makes our approach applicable to local infrastructure beyond drinking water, following a long tradition of estimating valuation for non-traded goods. Moreover, alternatives such as bottled water sales from consumer-panel scanner data cannot provide the fine-grained geographic and time variation and individual demographics we get from housing transactions.

[21]Relative to a hedonic approach inverting the equilibrium price surface, this allows us to recover preferences estimates that incorporate inframarginal house-buyers (Wong, 2018).

Note that having estimated the parameters for (13), which govern the demand for houses, we quantify resident i's utility $u_i(q_t, r_{it})$ as their indirect utility at the chosen house at t, measured in dollars by

$$u_i(q_t, r_{it}) = \underbrace{\frac{\theta_q(y_{it})}{\theta_p(y_{it})}}_{\equiv\omega(y_{it})} q_t - r_{it} + \iota_{it}$$

Estimation then proceeds in two stages, following (Bayer et al., 2007). In a first step, we estimate the house-specific utilities $\boldsymbol{\delta}$ and non-linear parameters $\boldsymbol{\theta_2}$ via maximum likelihood. The estimation sample in this first stage includes every residential property sold in California between 2002 and 2019. To minimize arbitrary restrictions on potential buyers' choice sets, we consider large geographic markets, as shown in Appendix Figure B4. In Appendix B, we provide details on the computational approach adopted to estimate demand with a large number of products per market.

Our second step recovers the parameters governing residents' average willingness-to-pay $\boldsymbol{\theta_1}$ by regressing the house-specific utilities $\boldsymbol{\delta}$ on house characteristics, prices, and water quality. Due to the endogeneity of house prices, we introduce an instrument for house prices. Our instrument consists of a reduced-form approximation to the optimal instrument.[22] To account for time-varying unobserved neighborhood attributes that influence water pollution and might be correlated with unobserved quality, we include boundary-year fixed effects, $\psi_{b_j t}$, which equals one if house j is located within 500m distance of a PWS service area boundary b in year t. Ideally, each boundary area lies within a single jurisdiction, allowing the fixed effect to capture factors such as school quality and local policies (e.g., zoning or crime prevention). However, since this condition might not always be met in the data, as a robustness, we additionally control for school fixed effects and city/county government fixed effects, which we identify using boundary areas that span multiple school districts or jurisdictions.

6.5. **Estimation of the policymaker's preferences.** Recall that we fix the discount factor β; normalize the direct payoff from not investing as zero; assume that the choice shocks ϵ_{st} are i.i.d. Type-I extreme value with unit scale; and have identified

[22]We intend to eventually compute the model-implied approximation to the optimal instrument, following Bayer et al. (2007); Calder-Wang (2021). As such, we see our current instrument as a way to get good starting point for $\tilde{\boldsymbol{\theta_1}}$ to compute the corresponding optimal instruments $\boldsymbol{p^{IV}}(\tilde{\boldsymbol{\theta_1}}, \hat{\boldsymbol{\theta_2}})$. This multi-step approach is proposed in Bayer and Timmins (2007). We provide more details in Appendix B.

the state transition. Therefore, by Hotz and Miller (1993), we identify the choice-specific value difference, which is the sum of the continuation value difference and the flow utility difference:

$$-\kappa(\boldsymbol{x}_t) \; + \int \lambda(\rho(\boldsymbol{x}_t))\, \tilde{\pi}(\boldsymbol{y}, \boldsymbol{x}_t) \left(\mathbb{E}\big[u(\boldsymbol{o}_t, \boldsymbol{y}) \mid 1, \boldsymbol{s}_t\big] - \mathbb{E}\big[u(\boldsymbol{o}_t, \boldsymbol{y}) \mid 0, \boldsymbol{s}_t\big] \right) dF_{\boldsymbol{y},t}(\boldsymbol{y}),$$

with $\boldsymbol{o}_t = (q_t, \bar{r}_t, q_{t-1}, \bar{r}_{t-1})$, water service outcomes. The flow utility difference is identified up to a location normalization because of the model's *renewal* feature, where investment resets $l_t = 0$, and an exclusion restriction, where transitory weather shocks (components of $\boldsymbol{x}_t$) affect election prospects (and thus current payoffs) but not state transitions, as in Rust (1987). Now, to separately identify $\kappa(\cdot)$ and $\lambda(\cdot)$, we exploit variation in the already-identified objects inside the integral and our exclusion restriction that resident preferences for water services do not depend on their jurisdiction's political features such as election timing. Policymaker preferences are then estimated by maximum likelihood, solving the dynamic program at each parameter guess with a nested fixed-point algorithm.

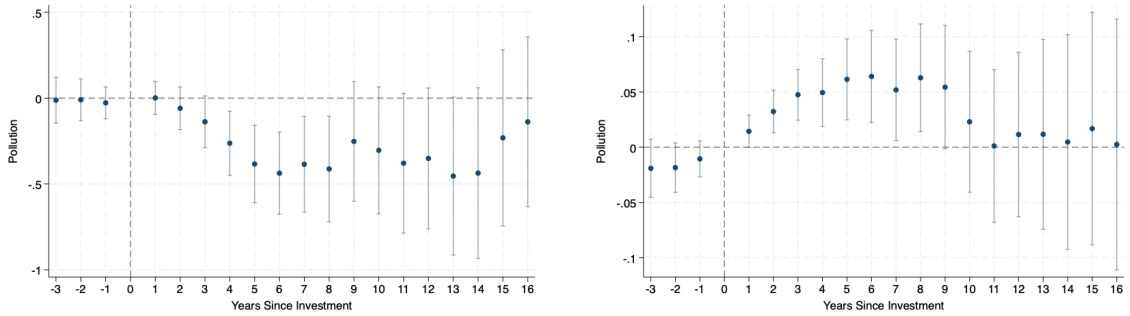
7. Estimation results

This section reports the empirical findings. First, we present the estimated water quality and cost dynamics, highlighting the effectiveness and heterogeneous cost of capital investments born by residents in drinking water infrastructure. Next, we turn to residents' preferences, summarized by their marginal willingness to pay for quality, $\omega(\boldsymbol{y})$. Together, these estimates allow us to gauge the magnitude of political distortions and to assess how they alter the effectiveness of standard policy incentives for investment.

7.1. Effectiveness and cost of investments.

Panel (A) of Figure 3 plots the event-study coefficients from (9). Relative to the pre-investment baseline, water-system upgrades produce a sharp and statistically significant decline in contamination: in years 3-8 after installation the log pollution index falls by about 0.27, a 27% reduction in drinking water pollution either the EPA violations or contaminant readings above the MCL. This magnitude closely matches the 0.5 pp drop (roughly 33 %) reported by Keiser et al. (2023) for systems receiving Drinking Water State Revolving Fund loans. Appendix C shows that the effect is fairly uniform across systems with different observable characteristics, although higher per-capita spending yields larger quality gains. Notably, the improvement erodes over time and has vanished in sixteen years.

FIGURE 3. Effects of Investment on Water Pollution and Rate



(A) Water pollution (B) Water rate

Notes: The figure displays the estimated coefficients for the dynamic effects of local government drinking water infrastructure investments on water pollution (Panel A) and average per-household water rates (Panel B), using data over the period 2003-2022. Water pollution is measured as the share of pollutants with readings that exceed the federal maximum contaminant level, including EPA health violations as a separate, additional pollutant.

Panel (B) of Figure 3 presents the event-study estimates from (11). Relative to the pre-investment baseline, locally financed upgrades lead to a statistically significant increase in water rates of about 5%, which persists for roughly nine years following the investment. In contrast, when investments are fully subsidized by federal or state funds, water rates remain unchanged. This stark difference underscores how the financing structure determines whether the costs of infrastructure are passed on to residents.

7.2. **Resident preferences for drinking water quality.** Table 4 presents the demand estimates specified in (12), which result from the two-stage procedure of Bayer et al. (2007). As expected, the non-linear coefficient on distance θ_d and the linear coefficient on price θ_{p0} are negative: residents dislike more expensive houses and value short commutes. In addition, the estimates on house characteristics also have the expected signs: residents value larger lots, houses with more bedrooms, and neighborhoods with a higher share of college graduates.

Finally, we estimate a negative, and significant, coefficient on water pollution, implying that residents do value cleaner drinking water. Our non-linear parameters imply that high-income buyers are not as pollution-sensitive as low-income residents, suggesting that they can compensate for higher pollution by reducing their exposure through expensive equipment like filtration systems. In line with the existing

TABLE 4. Residents Value Drinking Water Quality Heterogeneously

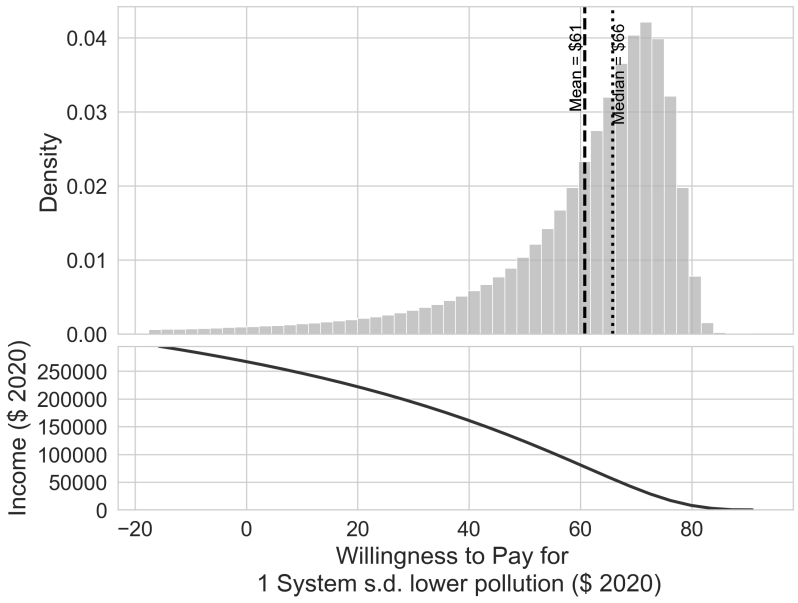
	Without Neigh. Pref.	With Neigh. Pref.
First Stage		
Income (log) × Water pollution	0.008 (0.550)	0.025 (0.115)
Income (log) × House price	36.241 (0.044)	32.116 (0.118)
Distance to workplace	-6.792 (0.040)	-6.799 (0.273)
Income (log) × Block-group income		0.480 (0.236)
Second Stage (IV) - Dep. Var: $\hat{\delta}$		
House user cost (price + rate)	-66.187 (1.257)	-59.742 (1.116)
Water pollution	-0.020 (0.007)	-0.023 (0.009)
Fraction of white	0.003 (0.000)	0.002 (0.000)
Fraction of college educated	0.968 (0.047)	0.933 (0.046)
# Bedrooms (log)	0.037 (0.007)	0.038 (0.006)
Sq. Footage (log)	1.066 (0.030)	0.971 (0.027)
Other house attributes	Yes	Yes
Border × Year FE, Market FE	Yes	Yes
Observations	870,637	870,637

Notes: This table presents the estimation results from the two-stage procedure of Bayer et al. (2007). Standard errors are in parentheses.

literature (Bayer et al., 2007, 2016), we also find that higher-income buyers are less price sensitive. Taken together, the two effects imply that higher-income residents display a higher marginal willingness to pay for water quality than lower-income residents. In Appendix B.2, we provide additional results that confirm that house prices are responsive to water pollution. First, we show hedonic regressions that confirm that residents dislike water pollution, including school and city fixed effects to limit variation coming from differences in jurisdictions besides water systems. Second, to ensure our results don't reflect unobserved amenity differences across boundaries, in Appendix B.2 we also present an event study evaluating the changes in house prices within an entity following a health-based drinking water violation.

Based on our estimates from column (1) of Table 4, we provide the distribution of the corresponding annualized willingness to pay for a one standard deviation decrease in water pollution across households, water systems, and periods, as measured in 2020 CPI-adjusted dollars, in Figure 4. We find that the average willingness to pay per year is $61. Combining or heterogeneous price and pollution sensitivities, once we account we account for heterogeneous sorting based on preferences for richer neighbors, we find that willingness to pay decreases with household income.

FIGURE 4. Residents' Willingness to Pay for Drinking Water Quality Improvement



Notes: This graph presents the distribution of the willingness to pay $\omega(\boldsymbol{y})$ for a one standard deviation increase in water quality across households and water systems per year, where the standard deviation is computed across water systems-year, in CPI-adjusted 2020 dollars, using the demand estimates from column (1) of Table 4. Each system in this graph is represented with an equal population.

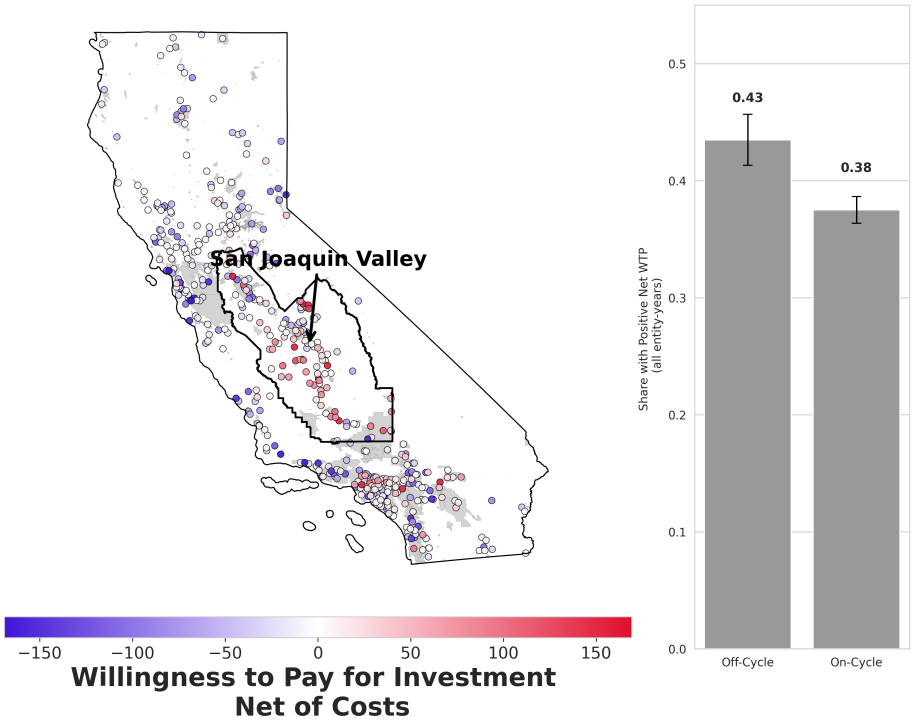
7.3. **Quantifying the net benefits of investment.** Taken together, our estimates for residents' preferences and investment costs can be used to quantify the value of investment in clean water infrastructure to residents. To this end, we compute the net change in the residents' utility $u(q_t, r_t; y_{it})$ following an investment in drinking water infrastructure:

$$\Delta u(q_t, r_t; y_{it}) = \omega_q(y_{it})\Delta q_t - \Delta r_t, \tag{14}$$

where the changes Δr_t and Δq_t are quantified in Sections 7.1, and the parameter $\omega_q(y_{it})$ captures the willingness to pay for clean water estimated in Section 7.2.

Figure 5 depicts the average net benefit of investment across utilities in our sample. We find that for 39% of utilities serving in total 65% of the population, residents would benefit from investment. Among these utilities, the average net benefit is $42 per year per household, and the benefits from investment are, on average, 4.5 times higher than the cost.

FIGURE 5. Net benefit of investment across space and election cycles



Notes: On the left panel, the figure displays the net investment benefits in 2020 USD as defined in equation (14) averaged across residents served by a community water system, represented by each polygon. On the right panel, the bar graph presents the fraction of cities whose total net benefit of an investment is positive, for the two groups cities depending on their mayoral and city council election schedule coinciding with the Election Days or the Presidential or state-wide primary dates ("on-cycle" vs. "off-cycle")

These results suggest that there is under-investment in water infrastructure for a sizable portion of the population in our sample. Can the lack of political accountability partially explain the under-investment in drinking water infrastructure? To answer this question, Figure 5 showcases the share of utilities for which we estimate positive net returns from investment, separately for utilities that hold mayoral and council member elections on federal or gubernatorial Election Days ("on-cycle") and those that do not ("off-cycle"). Under-investment is more prevalent in off-cycle utilities: 43% of these utilities would benefit from investment -about 5 percentage points higher than the corresponding share among on-cycle utilities. This result is consistent with the idea that lower electoral accountability in "off-cycle" cities creates a wedge between the local politician's investment decisions and its constituents' needs, leading to higher under-investment.

TABLE 5. Estimates of the Role of Political Factors in Investment

	Coefficient	Std. Error
$\boldsymbol{\kappa}$: **Capacity**		
Constant	3.276	(0.371)
Number of skilled employees (log)	-0.157	(0.028)
Small-sized	0.454	(0.039)
$\boldsymbol{\lambda}$: **Weight for constituent welfare**		
Constant	-15.874	(3.160)
Prob. of holding contested elections (estimated)	9.613	(3.349)
County FE	Yes	
Observations	8,986	
Avg. Neg. log-likelihood	0.230	

Notes: This table presents coefficient estimates for the parameters governing the role of local institutional and political factors. Coefficients are estimated using Nested Fixed Point (NFXP) maximum likelihood and standard errors are computed using the approximate Hessian at the parameter estimates. The estimation sample focuses on the 80% of entities with the largest average expected pollution, so that investment is more plausibly related to pollution abatement. The specification for both κ and λ assumes a log-linear functional form, with covariates entering additively in the exponent.

7.4. **How local political factors shape infrastructure investment decisions.** Table 5 presents the estimates from the dynamic investment model under political constraints presented in Section 5. We group parameters into two categories: those governing the private costs of investment borne by local officials, which we refer to as *capacity* parameters, and those governing the responsiveness of investment decisions to constituents preferences, which we refer to as *accountability* parameters.

Consistent with the evidence in Section 4, our estimates reveal that when elections are more likely to be contested, local officials are more responsive to their constituents welfare. As a result, we show in Appendix Figure D10 that our estimated model fits two important patterns from the data well: observed investment probability falls when predicted investment costs rise and increases with predicted pollution. Turning to capacity parameters, we find notably that the private cost of investments $\boldsymbol{\kappa}$ is higher for small systems and decreases with the number of highly paid employees working at the water system. Ignoring the effect that changing private costs would have on continuation values, increasing the skilled workforce at a water system by 10% would increase the log-odds of investment by about 1.6%. This result

echoes recent findings on the role of skilled employees in shaping the ability of local governments to invest in infrastructure (Liscow et al., 2025).[23]

8. Quantification and policy implications of accountability

We use our estimated model to quantify the role of local electoral accountability in shaping drinking infrastructure investments. We consider a counterfactual where governments face consistently contested elections throughout the sample period. Based on our estimates (Section 7), this increases the weight λ_s placed on voter welfare in government decision-making.[24] We first simulate the resulting investment paths and their implications for water pollution and rates. Next, we evaluate how the effects of existing federal and state subsidies on drinking water quality would differ under this higher electoral accountability scenario.

8.1. **Impact of accountability on infrastructure investment.**

Under the fully-contested counterfactual, we hold the generosity of drinking water subsidies constant and fix the impact of investment on water rates to our event study estimates (Figure 3), and we solve for the revised optimal investment decisions of local officials across the state space.[25] As we describe below, overall, our results showcase that greater electoral accountability increases investment in the systems that need it most.

Panel (A) of Figure 6 plots average investment probability under both the observed baseline and the fully contested counterfactual, split by quintiles of underlying system pollution. We define underlying pollution as $\frac{1}{T}\sum_t \alpha_{st}$ from the estimated pollution law of motion. Under the fully contested counterfactual, investment is reallocated toward systems with the greatest need and modestly reduced in cleaner systems relative to the baseline.[26] In the most pollution-prone systems (right panel),

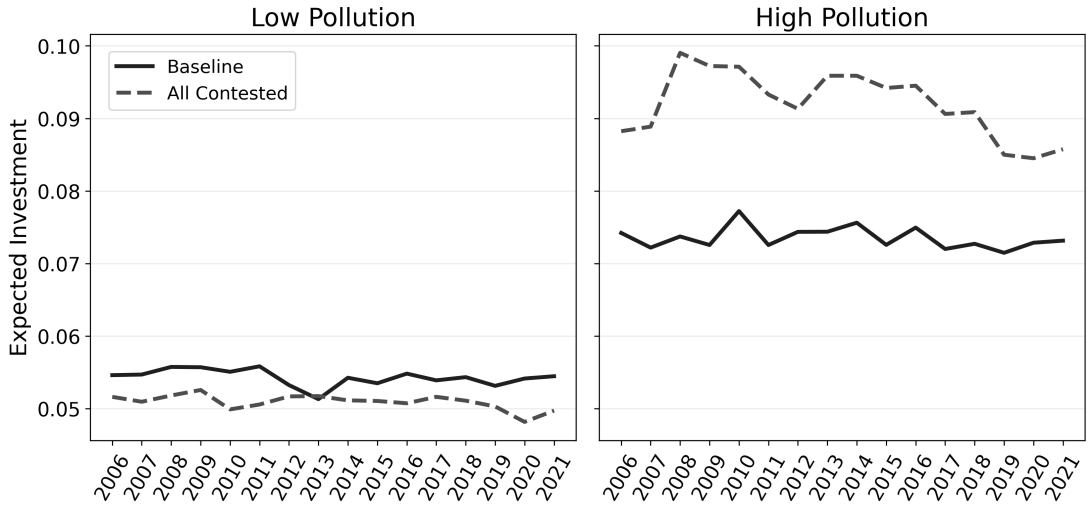
[23]Many of primary state and federal funding sources are restricted for capital improvements, and cannot be used to alleviate insufficient capacity due to the limited availability of skilled employees, as noted by this 2020 report from the California Legislative Analyst Office.

[24]Although such an increase in accountability may appear ambitious, recent policy changes illustrate that shifts in electoral competitiveness are feasible. For example, California Senate Bill 415 (the California Voter Participation Rights Act) induced many jurisdictions to move from off-cycle to on-cycle elections, which increased election contestedness (Figure A1 in the Appendix). This episode provides a concrete policy precedent for changes that enhance local electoral accountability.

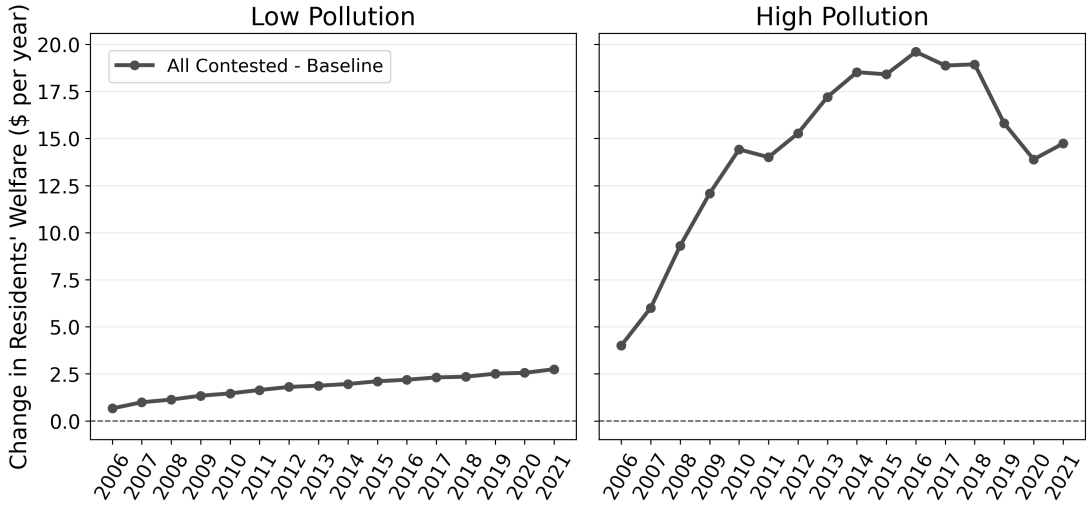
[25]Specifically, we initialize every system to their observed state in 2003 and simulate 1,000 paths of investment, pollution, and water rates until the end of our estimation period, 2022.

[26]This pattern closely aligns infrastructure decisions with the stated objectives of state and federal drinking water programs. See, for example, the California DWSRF Intended Use Plan, which notes that "The DWSRF program finances infrastructure improvements to mitigate drinking water risks and support the human right to water."

FIGURE 6. Role of Electoral Accountability



(A) Investment



(B) Resident welfare

Notes: Panel (A) plots the average investment probability under the counterfactual scenario in which all elections are contested, averaged across simulated paths and across governments grouped by quintiles of average expected pollution. For each government i, average expected pollution is defined as the sample mean of α_{st} in (10). The "Low pollution" group comprises governments with the lowest expected pollution, while "High pollution" comprises those with the highest. The solid line ("Baseline") shows observed investment probabilities, and the dashed line ("All contested") shows the simulated probabilities under the counterfactual. Panel (B) reports the average change in resident welfare under the counterfactual, averaged across simulated paths and governments.

investment probability rises by over 2 percentage point from a 7% baseline - an increase exceeding 25%. Given the estimated effectiveness of investment (Section 6), this translates into substantial pollution reductions of up to 0.35 standard deviations. By contrast, accountability modestly reduces investment in the least pollution-prone systems, where the marginal benefits of further pollution reduction fail to outweigh the associated increases in water rates.

Panel (B) of Figure 6 plots the resulting change in constituent welfare, which weighs their willingness to pay for improved water quality against the higher water rates required to finance investment. By increasing λ_s, greater electoral accountability aligns government decisions more closely with constituents preferences, and therefore generates welfare gains. The magnitude of these gains is economically meaningful: in the most polluted systems, average household welfare increases as high as $20 per year, exceeding the median per-household cost of investment (approximately $18.5). These estimates likely understate the true welfare effects of accountability. As discussed in Section 6, our preference parameter estimates may be conservative with respect to the health and economic costs of drinking water contamination.[27]

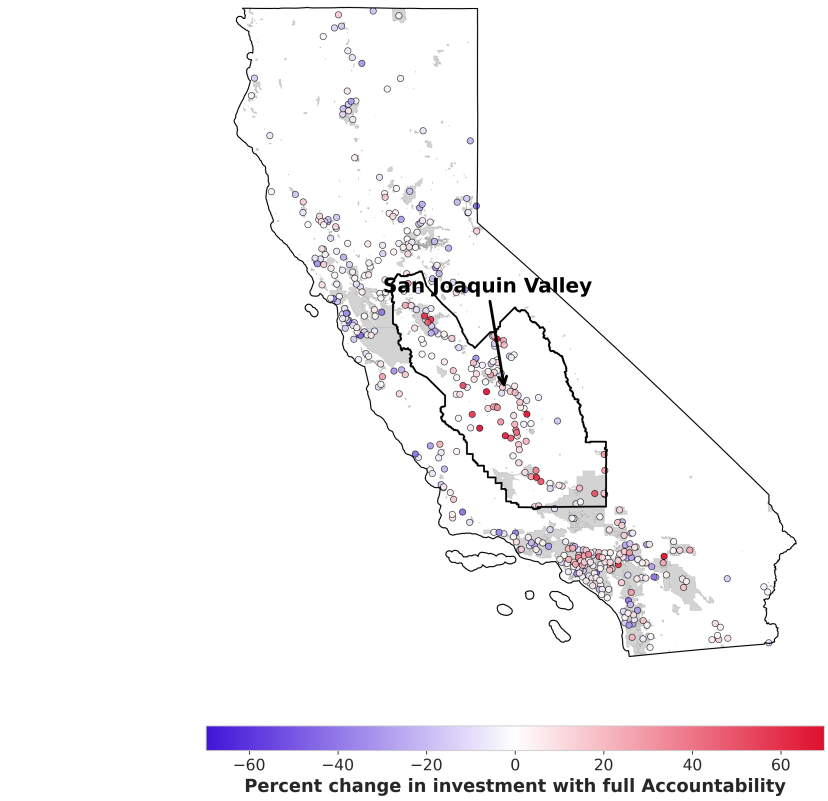
In addition to increasing the efficiency of infrastructure investment, greater electoral accountability has progressive distributional consequences. Figure 7(A) shows the spatial distribution of investment changes from moving to contested elections across California. Large increases in investments are concentrated in the San Joaquin Valley, an agricultural region in Central California facing persistent water quality issues (Balazs et al., 2011, 2012). Many of these systems also suffer from weak electoral competition at the local level.[28] As such, even when funding is available, the incentives for local officials in those water boards to address even pressing pollution issues may not be strong.[29] Consistent with this pattern, Figure 7(B) shows that accountability-induced investment increases are larger in more disadvantaged communities. Investment increases more in systems with lower share of college educated residents, higher shares of Hispanic residents, lower homeownership rates, and lower median household income. Thus, with greater accountability, investment patterns

[27]For example, Keiser et al. (2023) report that DWSRF loans and grants reduce mortality among older adults, and Hill and Ma (2022) documents adverse impacts of fracking-induced water pollution on infant health.

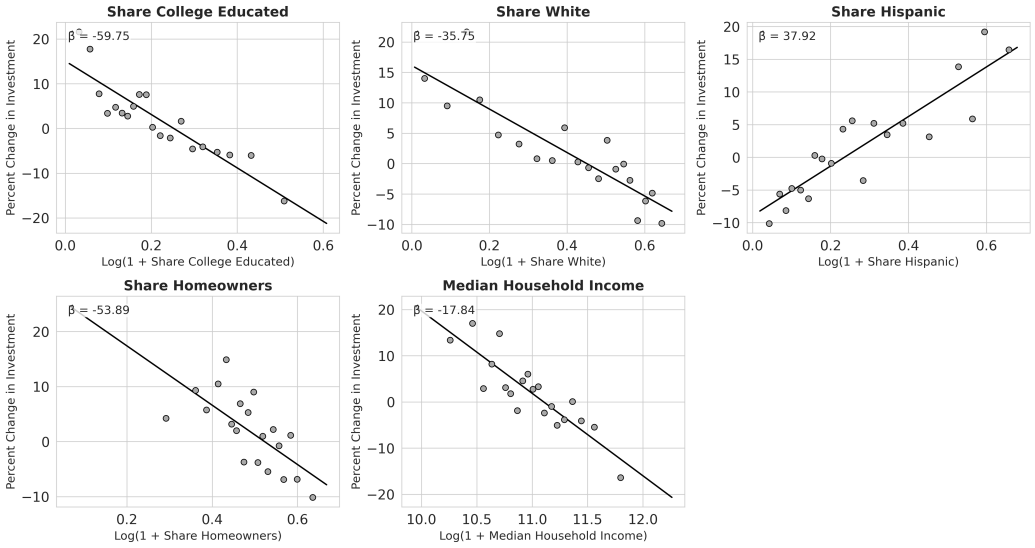
[28]A 2018 report by the non-profit Community Water Center notes that "In the local water boards of the southern San Joaquin Valley, contested elections are the exception, not the rule."

[29]Anecdotally, East Orosi CSD in the San Joaquin Valley accumulated 99 water quality violations over our sample period without undertaking any investment or holding a contested election; the system was placed under a state administrative order in 2022.

FIGURE 7. Heterogeneous Effects of Accountability on Investment



(A) Across Space



(B) Across Demographics

would align more closely with state and federal policy objectives of improving water quality especially in small and disadvantaged communities.

8.2. **Impact of accountability on the returns to public funding.** We next investigate the impact of limited accountability on the cost-effectiveness of federal and state funding policies for drinking water infrastructure. A vast web of state and federal funding programs ensures that only a fraction of infrastructure improvement costs are passed on to residents in the form of water rates increases. At the helm of these initiatives, Drinking Water State Revolving Funds (DWSRF) have allocated more than $40 billion in loans since 1996 (EPA, 2019), but these funds may not necessarily flow to the systems that need them most (Hansen et al., 2021).

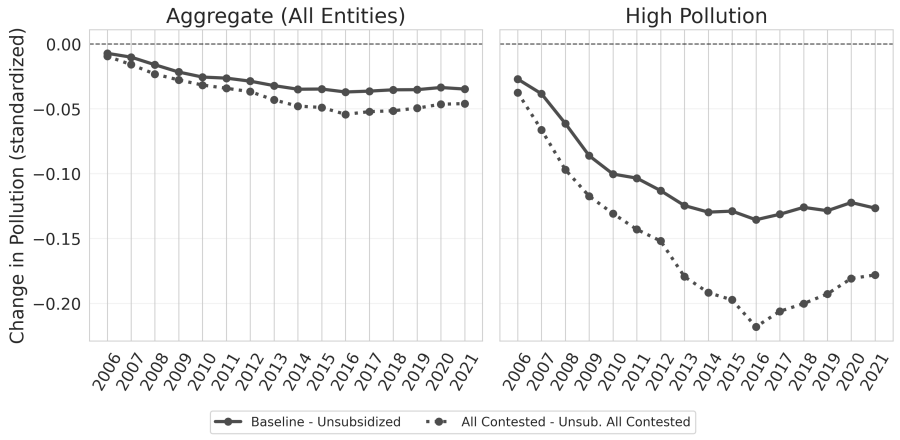
We thus run a second counterfactual where we quantify how limited accountability undermines the targeting of state and federal funding programs. First, we compute the counterfactual water rate increases that would be needed to fund infrastructure investment in the absence of external subsidies.[30] Then, using our estimated model, we solve for the revised optimal investment decisions of local officials across the state space, under both the fully-contested environment, and the observed one.

Figure 8(A) shows that current federal and state subsidies decrease pollution by around 0.028 standard deviations on average and up to 0.105 standard deviations for the highly polluted areas. The decrease would be 37% larger on average in the fully-contested scenario, and 47% larger for the most polluted systems. Figure 8(B) illustrates the mechanism by plotting subsidy take-up—measured in dollars per resident—against the (gross) welfare gains from investment in both scenarios. In the baseline, subsidy take-up is relatively higher when welfare gains are low and relatively lower when welfare gains are high, compared to the fully-contested benchmark. This pattern reflects a wedge created by limited accountability between politicians' investment decisions and residents' welfare gains, leading to a misallocation of subsidies toward investments that are less effective at reducing pollution.

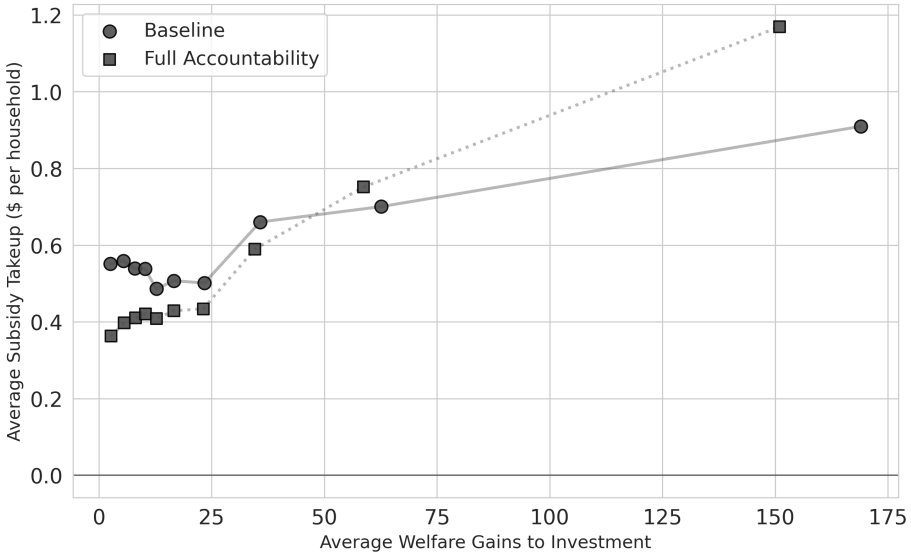
These results imply that alleviating fiscal constraints alone may not ensure efficient decentralized infrastructure provision. When local accountability is limited, subsidies may fail to flow to the highest-return projects. Our findings therefore call for approaches to local infrastructure funding that strengthen accountability or mitigate local discretionary distortions, for example through administrative oversight or

[30]See the detailed construction of the no-subsidy water rate increases in the Appendix.

FIGURE 8. Effects of Infrastructure Subsidies by Accountability



(A) Impact on pollution



(B) Subsidy take-up by welfare benefits

Notes: Panel (A) shows the effect of federal and state subsidies on drinking water pollution under both the baseline and the fully-contested scenarios, while holding the subsidy environment constant. "Baseline - Unsubsidized" (solid line) indicates that subsidy effectiveness is simulated at observed levels of election contestedness. "All Contested - Unsub. All Contested" (dashed line) indicates a simulation under a scenario where every election has at least two contenders. Panel (B) shows the average subsidy take-up per household in 2020 USD as a function of the "gross" welfare gains from investment, that is, the welfare benefits that would accrue to residents only from pollution decreases, ignoring associated water rate increases, expressed in 2020 USD per household, per year.

institutional support that helps local governments to identify, plan and access central government funding opportunities.

9. Conclusion

We provide evidence that infrastructure investment decisions made by local governing bodies may not align with those that maximize constituent welfare. Using a novel data set on investment and performance of drinking water systems in California, we document that weaker electoral competition is associated with worse pollution outcomes and lower take-up of state and federal subsidies. Even within systems that transition to more competitive election cycles ("on-cycle"), we find that pollution falls over time. We then quantify the welfare consequences of higher pollution and water rates using a structural housing choice model, and measure the distortions generated by local political frictions using a dynamic investment framework. These frictions are costly: if all local elections were contested, with at least two candidates competing for every seat, investment would increase, pollution would fall, and welfare would improve, particularly in struggling systems serving disadvantaged communities.

More broadly, our findings highlight the role of political accountability in shaping allocative efficiency within decentralized public investment. Fiscal resources alone do not ensure efficient outcomes; the translation of transfers into welfare gains depends on the incentives faced by local policymakers to align their investment decisions with constituent welfare. Reforms that strengthen electoral competition or enhance accountability may therefore complement financial support by improving both the level and the targeting of investment. Designing infrastructure policy without accounting for political incentives risks limiting the impact of large-scale public spending. Thus, an important direction for future research is to examine which institutional reforms—such as electoral design, information provision, or oversight mechanisms—most effectively improve accountability, and how federal or state policy can be structured given the institutional constraints local governments face.

References

Adelino, Manuel, Igor Cunha, and Miguel A Ferreira, "The economic effects of public financing: Evidence from municipal bond ratings recalibration," *The Review of Financial Studies*, 2017, *30* (9), 3223–3268.

Agrawal, Ashwini K and Daniel Kim, "Municipal bond insurance and public infrastructure: Evidence from drinking water," *Available at SSRN 3813348*, 2022.

Allaire, Maura, Haowei Wu, and Upmanu Lall, "National trends in drinking water quality violations," *Proceedings of the National Academy of Sciences*, 2018, *115* (9), 2078–2083.

Anzia, Sarah F, "Election timing and the electoral influence of interest groups," *The Journal of Politics*, 2011, *73* (2), 412–427.

Aruoba, S Borağan, Allan Drazen, and Razvan Vlaicu, "A structural model of electoral accountability," *International Economic Review*, 2019, *60* (2), 517–545.

Avis, Eric, Claudio Ferraz, and Frederico Finan, "Do government audits reduce corruption? Estimating the impacts of exposing corrupt politicians," *Journal of Political Economy*, 2018, *126* (5), 1912–1964.

Bai, Shaojie, J Zico Kolter, and Vladlen Koltun, "Deep equilibrium models," *Advances in neural information processing systems*, 2019, *32*.

Balazs, Carolina L, Rachel Morello-Frosch, Alan E Hubbard, and Isha Ray, "Environmental justice implications of arsenic contamination in California's San Joaquin Valley: a cross-sectional, cluster-design examining exposure and compliance in community drinking water systems," *Environmental Health*, 2012, *11* (1), 84.

Balazs, Carolina, Rachel Morello-Frosch, Alan Hubbard, and Isha Ray, "Social disparities in nitrate-contaminated drinking water in California's San Joaquin Valley," *Environmental health perspectives*, 2011, *119* (9), 1272.

Bayer, Patrick and Christopher Timmins, "Estimating equilibrium models of sorting across locations," *The Economic Journal*, 2007, *117* (518), 353–374.

_____, Fernando Ferreira, and Robert McMillan, "A unified framework for measuring preferences for schools and neighborhoods," *Journal of political economy*, 2007, *115* (4), 588–638.

_____, Robert McMillan, Alvin Murphy, and Christopher Timmins, "A dynamic model of demand for houses and neighborhoods," *Econometrica*, 2016, *84* (3), 893–942.

Bennear, Lori S and Sheila M Olmstead, "The impacts of the "right to know"': Information disclosure and the violation of drinking water standards," *Journal of Environmental Economics and Management*, 2008, *56* (2), 117–130.

Berry, Christopher R, *Imperfect union: Representation and taxation in multilevel governments*, Cambridge University Press, 2009.

Besley, Timothy and Anne Case, "Does electoral accountability affect economic policy choices? Evidence from gubernatorial term limits," *The Quarterly Journal*

of Economics, 1995, *110* (3), 769–798.

Biasi, Barbara, Julien Lafortune, and David Schönholzer, "What works and for whom? Effectiveness and efficiency of school capital investments across the US," *The Quarterly Journal of Economics*, 2025, *140* (3), 2329–2379.

Black, Sandra E, "Do better schools matter? Parental valuation of elementary education," *The quarterly journal of economics*, 1999, *114* (2), 577–599.

Bordeu, Olivia, "Commuting infrastructure in fragmented cities," *Job Market Paper, University of Chicago Booth School of Business*, 2023, *6*.

Brancaccio, Giulia, Myrto Kalouptsidi, and Theodore Papageorgiou, "Investment in infrastructure and trade: The case of ports," Technical Report, National Bureau of Economic Research 2024.

Burgess, Robin, Remi Jedwab, Edward Miguel, Ameet Morjaria, and Gerard Padró i Miquel, "The value of democracy: evidence from road building in Kenya," *American Economic Review*, 2015, *105* (6), 1817–1851.

Calder-Wang, Sophie, "The distributional impact of the sharing economy on the housing market," *Available at SSRN 3908062*, 2021.

Callaway, Brantly and Pedro HC Sant'Anna, "Difference-in-differences with multiple time periods," *Journal of econometrics*, 2021, *225* (2), 200–230.

Calonico, Sebastian, Matias D Cattaneo, and Rocio Titiunik, "Robust nonparametric confidence intervals for regression-discontinuity designs," *Econometrica*, 2014, *82* (6), 2295–2326.

Cellini, Stephanie Riegg, Fernando Ferreira, and Jesse Rothstein, "The value of school facility investments: Evidence from a dynamic regression discontinuity design," *The Quarterly Journal of Economics*, 2010, *125* (1), 215–261.

Cengiz, Doruk, Arindrajit Dube, Attila Lindner, and Ben Zipperer, "The effect of minimum wages on low-wage jobs," *The Quarterly Journal of Economics*, 2019, *134* (3), 1405–1454.

Conlon, Christopher and Jeff Gortmaker, "Best practices for differentiated products demand estimation with pyblp," *The RAND Journal of Economics*, 2020, *51* (4), 1108–1161.

de Benedictis-Kessner, Justin, "Off-cycle and out of office: Election timing and the incumbency advantage," *The Journal of Politics*, 2018, *80* (1), 119–132.

DiSalvo, Richard W and Elaine L Hill, "Drinking water contaminant concentrations and birth outcomes," *Journal of Policy Analysis and Management*, 2024, *43* (2), 368–399.

Dobbin, Kristin, Amanda Fencl, and Justin McBride, "California Community Water System institutional type update dataset," `https://doi.org/10.25338/B8KP92` 2023.

EPA, "A tractable approach to compare the hedonic and discrete choice frameworks," *Annual Report*, 2019.

Fajgelbaum, Pablo D, Cecile Gaubert, Nicole Gorton, Eduardo Morales, and Edouard Schaal, "Political Preferences and the Spatial Distribution of Infrastructure: Evidence from California's High-Speed Rail," Technical Report, National Bureau of Economic Research 2023.

Gandhi, Amit and Jean-François Houde, "Measuring substitution patterns in differentiated-products industries," *NBER Working paper*, 2019, (w26375).

Gowrisankaran, Gautam, Ashley Langer, and Mar Reguant, "Energy transitions in regulated markets," Technical Report, National Bureau of Economic Research 2024.

Hajnal, Zoltan L and Paul G Lewis, "Municipal institutions and voter turnout in local elections," *Urban Affairs Review*, 2003, *38* (5), 645–668.

_____, Vladimir Kogan, and G Agustin Markarian, "Who votes: City election timing and voter composition," *American Political Science Review*, 2022, *116* (1), 374–383.

Hansen, Katy and Megan Mullin, "Barriers to water infrastructure investment: Findings from a survey of US local elected officials," *PLoS Water*, 2022, *1* (8), e0000039.

_____, Sara Hughes, Andrea Paine, and James Polidori, *Drinking water equity: Analysis and recommendations for the allocation of the state revolving funds*, Environmental Policy Innovation Center, 2021.

Healy, Andrew and Neil Malhotra, "Myopic voters and natural disaster policy," *American Political Science Review*, 2009, *103* (3), 387–406.

Hill, Elaine L and Lala Ma, "Drinking water, fracking, and infant health," *Journal of Health Economics*, 2022, *82*, 102595.

Hotz, V Joseph and Robert A Miller, "Conditional choice probabilities and the estimation of dynamic models," *The Review of Economic Studies*, 1993, *60* (3), 497–529.

Iaryczower, Matias, Gabriel Lopez-Moctezuma, and Adam Meirowitz, "Career concerns and the dynamics of electoral accountability," *American Journal of Political Science*, 2024, *68* (2), 696–713.

Kashner, Zane, "Building with Externalities: Local Governments and Wind Farms," Technical Report, Working Paper 2026.

Keiser, David A, Bhashkar Mazumder, David Molitor, and Joseph S Shapiro, "Water Works: Causes and Consequences of Safe Drinking Water in America," 2023.

Lim, Claire SH and Ali Yurukoglu, "Dynamic natural monopoly regulation: Time inconsistency, moral hazard, and political environments," *Journal of Political Economy*, 2018, *126* (1), 263–312.

Liscow, Zachary, Cailin Slattery, and Will Nober, "State Capacity and Infrastructure Costs," *Yale Law & Econ. Research Paper*, 2025.

List, John A and Daniel M Sturm, "How elections matter: Theory and evidence from environmental policy," *The Quarterly Journal of Economics*, 2006, *121* (4), 1249–1281.

Madera, "Water Rate Study," `https://www.madera.gov/wp-content/uploads/2022/04/Madera_Water-Rate-Study-Report_FINAL2.pdf` 2022. [Online; accessed 15-January-2025].

Magnac, Thierry and David Thesmar, "Identifying dynamic discrete decision processes," *Econometrica*, 2002, *70* (2), 801–816.

Mian, Atif, Amir Sufi, and Francesco Trebbi, "The political economy of the US mortgage default crisis," *American Economic Review*, 2010, *100* (5), 1967–1998.

Muehlenbachs, Lucija, Elisheba Spiller, and Christopher Timmins, "The housing market impacts of shale gas development," *American Economic Review*, 2015, *105* (12), 3633–3659.

Oates, Wallace E, "Fiscal federalism," *New York*, 1972, *1.*

Ouazad, Amine and Matthew E Kahn, "Time Consistent Infrastructure Investments: Optimal Flood Protection Policies in Spatial Equilibrium," Technical Report, National Bureau of Economic Research 2025.

Paszke, A, "Pytorch: An imperative style, high-performance deep learning library," *arXiv preprint arXiv:1912.01703*, 2019.

Posenau, Kelly, "Debt and water: Effects of bondholder protections on public goods," Technical Report, Technical report, Working Paper 2021.

Rust, John, "Optimal replacement of GMC bus engines: An empirical model of Harold Zurcher," *Econometrica*, 1987, pp. 999–1033.

Sant'Anna, Pedro HC and Jun Zhao, "Doubly robust difference-in-differences estimators," *Journal of econometrics*, 2020, *219* (1), 101–122.

Sieg, Holger and Chamna Yoon, "Estimating dynamic games of electoral competition to evaluate term limits in us gubernatorial elections," *American Economic Review*, 2017, *107* (7), 1824–1857.

_____ **and** _____, "Electoral Accountability and Control in US Cities," *Journal of Political Economy*, 2022, *130* (11), 2985–3023.

Sileo, Gretchen, "Proactive and reactive infrastructure investment," *Available at SSRN 4274528*, 2023.

Timmins, Christopher, "Measuring the dynamic efficiency costs of regulators' preferences: Municipal water utilities in the arid west," *Econometrica*, 2002, *70* (2), 603–629.

Wong, Maisy, "A tractable approach to compare the hedonic and discrete choice frameworks," *Journal of Housing Economics*, 2018, *41*, 135–141.

Appendix A. Institutions

Figure A1. Effects of SB 415 (2015), California Voter Participation Rights Act

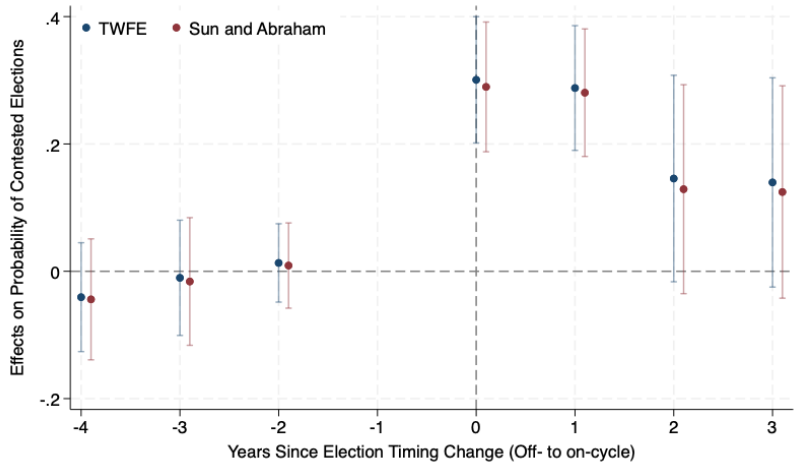
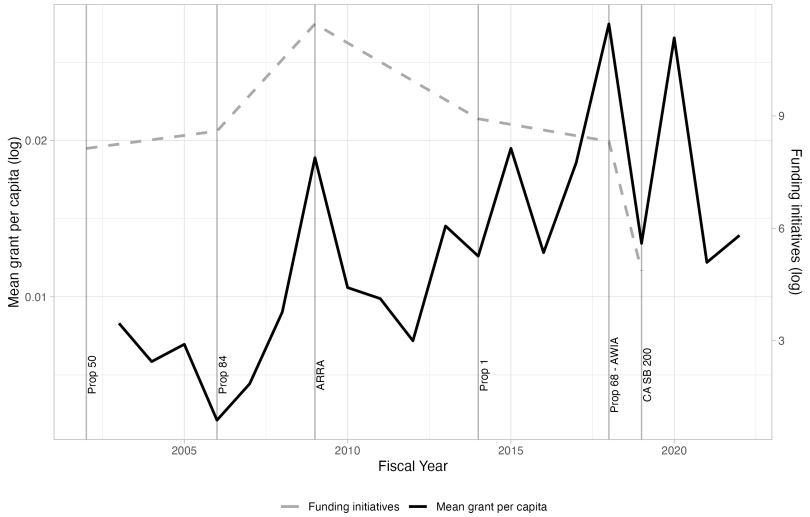
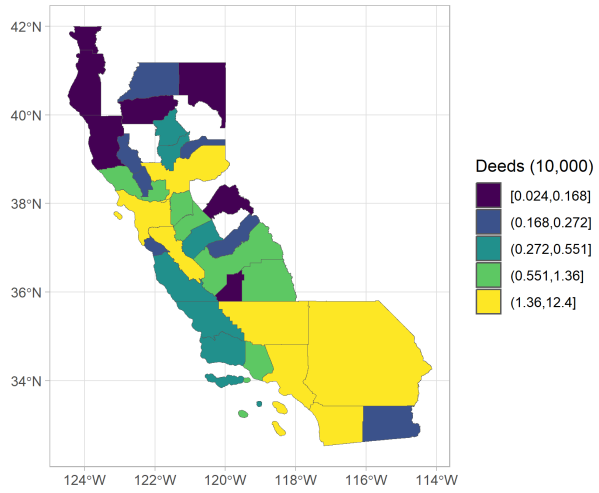


Figure A2. Statutory Funding Shocks and Local Drinking Water Grants



Notes: This figure plots the mean grant amount per capita (solid line, left axis, log scale) received by local governments in our sample for drinking water infrastructure investment over the study period. The dashed line (right axis, log scale) indicates the total pledged funding amounts from major federal/state statutes and state general obligation bonds, including Proposition 50 (2002), Proposition 84 (2006), Proposition 1 (2014), and Proposition 68 (2018). Vertical lines mark the enactment years of these statutes.

FIGURE B3. Geographic Markets used in Demand



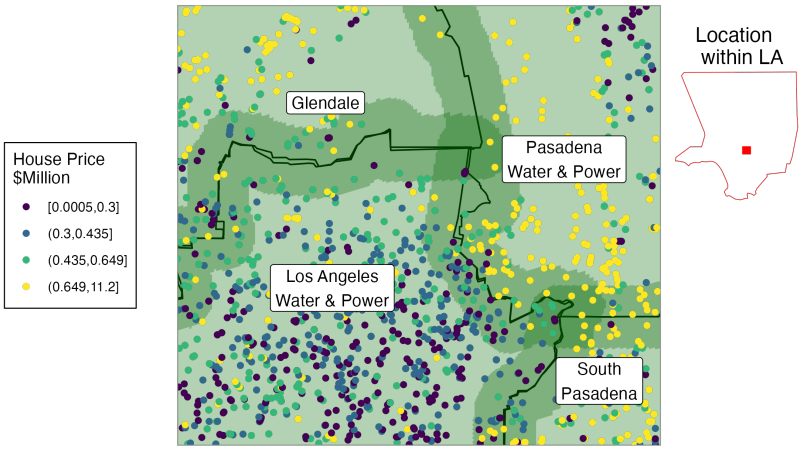
APPENDIX B. DEMAND ESTIMATION

B.1. **Estimation approach.**

B.1.1. *Data construction.* We follow Bayer et al. (2016) and link the CoreLogic housing transaction (deed) dataset to individual mortgage applicant data made available by the Home Mortgage Disclosure Act (HMDA). This results in a dataset of severeal millions house transactions between years 2003 and 2019 where we observe buyer income, ethnicity and race, and gender. In addition, we leverage the commuting matrix from the Longitudinal Household Employment Dynamics (LEHD) data to simulate a workplace for each house buyer based on the location of the house purchased. The simulated workplaces are therefore consistent with the empirical distribution of tract-to-tract commuting patterns in California in 2010. We then allocate these transactions to separate geographic housing markets defined using core-based statistical areas (CBSAs). Figure B4 below shows the resulting markets.

B.1.2. *Computational details.* As shown in Figure B4, the largest number of transactions in the geographic markets considered is above $100,000$. For the largest markets such as Los Angeles and the Bay Area, we define markets as geography–quarter instead of geography–year. This reduces the number of products but the largest market still consists in more than $30,000$ different transactions. As a result, we estimate our demand model using pytorch (Paszke, 2019). Our implementation follows the best

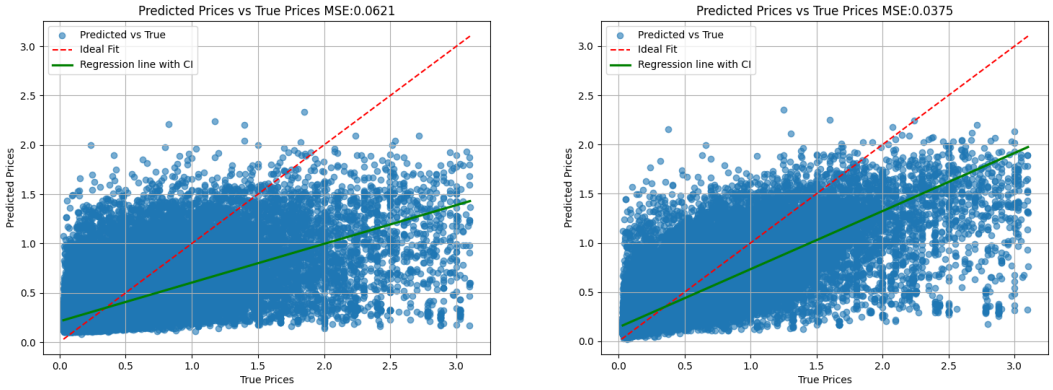
FIGURE B4. Boundary Discontinuity Design



practices Conlon and Gortmaker (2020), and we rely on pytorch for GPU acceleration and automatic differentiation. To efficiently apply automatic differentiation of the BLP contraction mapping, we follow the approach used in deep implicit layers models (see for instance, Bai et al. (2019)) and approximate the gradient by using an auxiliary fixed-point mapping that only requires the vector-Jacobian product which is directly provided by automatic differentiation libraries such as pytorch (Paszke, 2019). In doing so, we avoid both inverting a large $J \times J$ matrix and storing in memory all the intermediate tensors used to solve the contraction mapping at every parameter guess. With these changes, we use the L-BFGS optimization routine to estimate the parameters $\boldsymbol{\theta_2}$. We can estimate the demand model described in Section 6 in around five hours on a A100 GPU.

B.1.3. *Instrumental variable.* Following Bayer et al. (2007); Calder-Wang (2021) we intend to compute an instrument for prices using a guess of parameters and the model to construct the predicted house prices that result from variation in exogenous house characteristics only. To obtain an initial guess of the linear parameters $\boldsymbol{\theta_1}$, we need a preliminary instrument for house prices. We follow Gandhi and Houde (2019) and construct differences in exogenous house characteristics $\boldsymbol{x_h}$–number of bedrooms, bathrooms, building and land square footage, year built– where for a given characteristic ℓ we compute $z^1_{j\ell} = \sum_{k,1km \leq d_{hk} \leq 5km}(x_{j\ell} - x_{k\ell})$ and $z^2_{j\ell} = \sum_{k,1km \leq d_{hk} \leq 5km}(x_{j\ell} - x_{k\ell})^2$. Finally, we compute the best predictor of house prices, $\hat{p}_{jt} = \mathbb{E}[p_{jt}|\boldsymbol{x_{jt}}, \boldsymbol{z^1_{jt}}, \boldsymbol{z^2_{jt}}]$ using gradient boosted trees and cross-fitting. We show in Figure B5 that this instrument

FIGURE B5. Best Predictor of House Price



(A) $\mathbb{E}[p_{ht}|\boldsymbol{x_{ht}}]$ (B) $\mathbb{E}[p_{ht}|\boldsymbol{x_{ht}}, \boldsymbol{z^1_{ht}}, \boldsymbol{z^2_{ht}}]$

has a relevant first stage, and that information on houses around house j, excluding the ones directly in the vicinity of j, provides predictive power over and above the characteristics of house j itself.

B.2. **Additional hedonic regression results.** In addition to the structural demand model presented in the main text, we also estimate the following hedonic regression:

$$\log(p_{jt} + r_{jt}) = \beta q_{jt} + \gamma \boldsymbol{x}_{jt} + \psi_{bt} + \rho_s + \mu_g + \epsilon_{jt}, \tag{15}$$

where q_{ht} corresponds to different measures of water quality and where we control for house characteristics and neighborhood demographics, represented by $\boldsymbol{x}_{jt}$. To account for time-varying unobserved neighborhood attributes that influence house prices, we include boundary-year fixed effects, $\psi_{b_j t}$, which equals one if house h is located within 500m distance of a PWS service area boundary b in year t, ψ_{bt}. Ideally, each boundary area lies within a single jurisdiction, allowing ψ_{bt} to capture factors such as school quality and local policies (e.g., zoning or crime prevention). However, since this condition might not always be met in the data, we include boundary areas that span multiple school districts or jurisdictions. For this reason, we additionally control for school fixed effects and city/county government fixed effects.

Columns (1) and (2) of Table B1 show that water pollution, as measured by the average percentage by which contamination exceeds the federal standards across all pollutants during the calendar year of the sale, significantly impacts house prices. Across the specifications, a one standard deviation increase in water pollution is

TABLE B1. Do Residents Care about Water Quality?

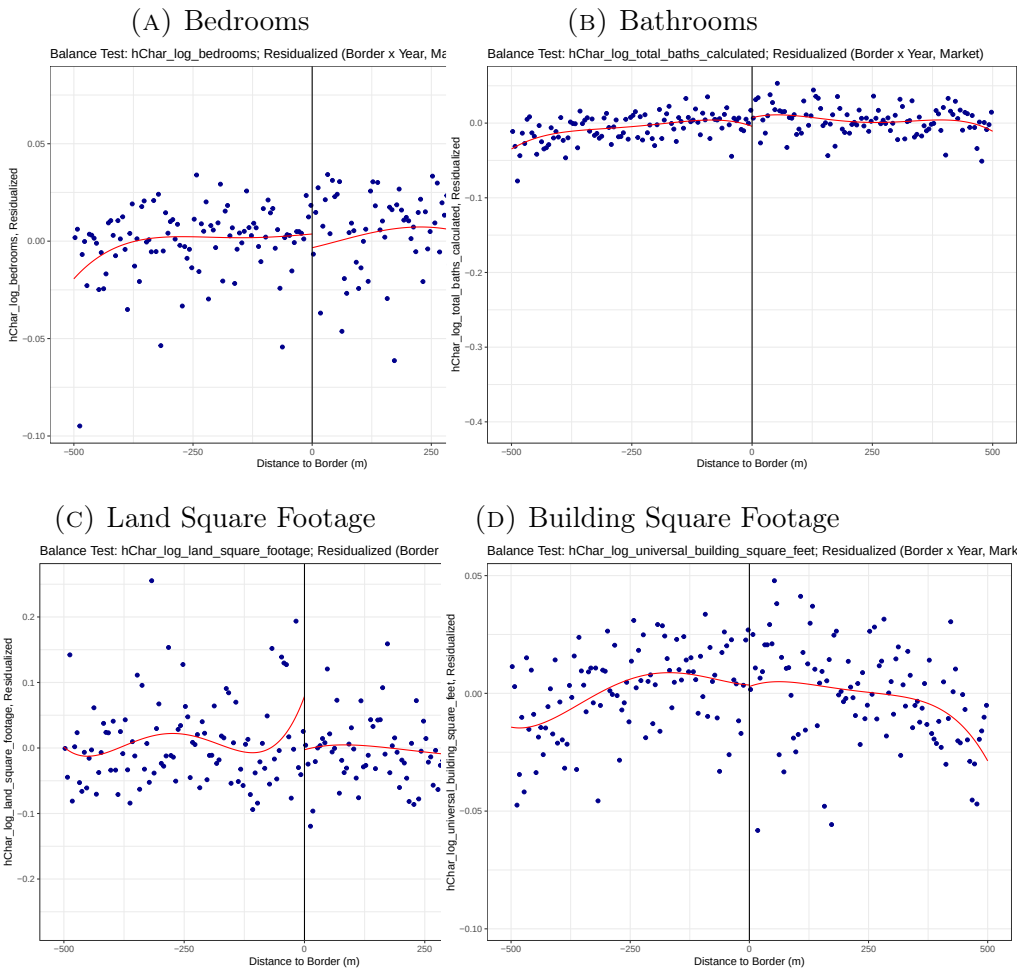
	House price + water expenses (log)		
	(1)	(2)	(3)
Water pollution[a]	-0.0117**	-0.0115**	-0.0135**
	(0.0053)	(0.0046)	(0.0054)
Any Tier I violation notification			-0.0349**
			(0.0103)
House and demographic attributes[b]	Yes	Yes	Yes
Border × Year FE	Yes	Yes	Yes
School district FE	No	Yes	Yes
City/county FE	No	Yes	No
Observations	910,209	910,209	910,226
R^2	0.805	0.810	0.805

Notes. This table presents OLS estimates based on house transactions in California from 2003 to 2019. Standard errors, clustered at the boundary-year level, are shown in parentheses. Statistical significance is indicated as follows: $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$. The dependent variable is the sum of the transaction price of the house and the present value of the water expenses at the year of sale, calculated over the next 14 years. a. "Water pollution" is measured as the average percentage of contamination exceeding the maximum contamination level (MCL), excluding disinfection byproducts and residualizing to correct for seasonality in contaminant readings. b. Controls include house attributes such as number of bedrooms, bathrooms, building and land square footage, and flexible controls for the year in which the house was built, along with Census block-group level demographics, including race, education, and median household income.

associated with a 1.4-1.6% decline in house prices. This magnitude aligns with the estimated effects of ozone pollution on house prices reported by Bayer et al. (2016).

Notably, Column (3) shows that, even when controlling for the continuous measure of water quality, a Tier I violation notification within the five years preceding the sale negatively impacts house prices, leading to an average 3.5% decrease. While this estimate is smaller than the 9.9-16.5% reduction attributed to nearby shale gas development (which heightens groundwater contamination risks) in Pennsylvania, as documented by Muehlenbachs et al. (2015), it remains both statistically and economically significant. This finding suggests that notifications may influence residents' incomplete perceptions of water quality or that residents prioritize short-term, acute water quality concerns linked to Tier I violations over longer-term issues that do not require immediate action. Following Bayer et al. (2007), we show on figure B6 that house characteristics do not jump discontinuously at the boundaries of public drinking water systems, lending further credibility to the boundary discontinuity design used in our main analysis.

FIGURE B6. House Attributes are Continuous across System Boundaries



Notes: We report regression discontinuity plots at the boundaries of water systems using the rdplot command from the R package rdrobust by Calonico et al. (2014). A negative distance to the boundary denotes a house on the "polluted" side of a boundary, and vice versa.

B.3. **Evidence from a within-jurisdiction design.** To alleviate concerns that the effects of water pollution on house prices reported from our reduced-form and structural water-boundary design exercises might stem form unobserved differences in amenities across public water systems boundaries, we also report estimates from an event study evaluating the changes in house prices within an entity following a health-based drinking water violation. Denoting a water system by s with violation

event h in year t, and a house by j, we estimate

$$\log p_{jsht} = \beta x_{jt} + \phi_{sh} + \mu_t \tag{16}$$

$$+ \sum_{\tau} \delta_{\tau}^{I} \; I_{s,t,h}^{\tau} \mathbb{1}\{h \text{ is Tier I}\} \tag{17}$$

$$+ \sum_{\tau} \delta_{\tau}^{II} \; I_{s,t,h}^{\tau} \mathbb{1}\{h \text{ is Tier II}\} + \epsilon_{jsht} \tag{18}$$

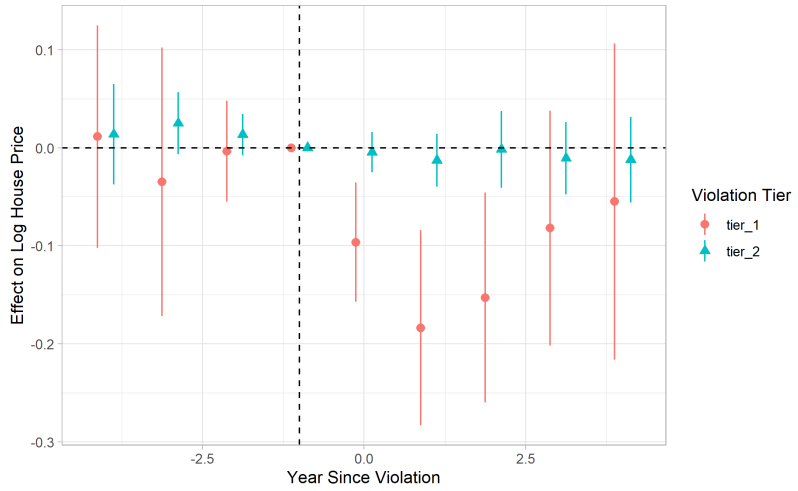
The treatment dummy $I_{s,t,h}^{\tau}$ equals 1 if the violation h occurred τ years from year t by water system s, and the specification includes event-specific water system fixed effects and year fixed effects, to account for systems with multiple violations over our sample. We control for a rich set of house characteristics x_{jsht}, as in our boundary discontinuity designs. We estimate the dynamic treatment effects δ_{τ} separately for Tier I and Tier II violations, to allow for different levels pollution to impact house prices differently. Moreover, the information available to residents and potential house buyers might be different as Tier I violations have stringent and immediate reporting requirements -residents are notified within 24 hours. We report dynamic treatment effects estimates in Figure B7. Consistent with Table B1 and with our main analysis, we find that Tier I violations adversely impact house prices in the medium run. As these effects are within water systems, they cannot be explained by cross jurisdiction differences in amenities.

Appendix C. Infrastructure Investments

Figure C8 presents the dynamic effects of investment on water pollution, using specification (9), for different sub-samples of the data: below-25th percentile and above-75th percentile groups for a given criterion (investment size, system size, and climate) or thee groups depending on the source of water. We do not find evidence for a large heterogeneity in the effects by the amount of the investment (Panel A), the size of the system as measured by the number of connections (Panel B), the source of water (groundwater, surface water drawn within the jurisdiction, or purchased surface water; Panel C), or the climate as represented by the average annual amount of rain during the period of study (Panel D).

Figure C9 describes the heterogeneity in the investment amount and the coupon rates. Panel (A) presents the distribution of the log amount of investment in 2020 CPI-adjusted dollars, based on the observed 1,979 investment cases. The median investment size is \$4.38 million while the average is \$23.64 million. The panel also

FIGURE B7. Event Study: Effects of Violations on House Prices



Notes: This figure presents estimates from the event-study specification 18estimated with a two-way fixed effects estimator. We report point estimates and 95% confidence intervals for Tier I and Tier II violations. Standard errors are clustered at the system-violation event level.

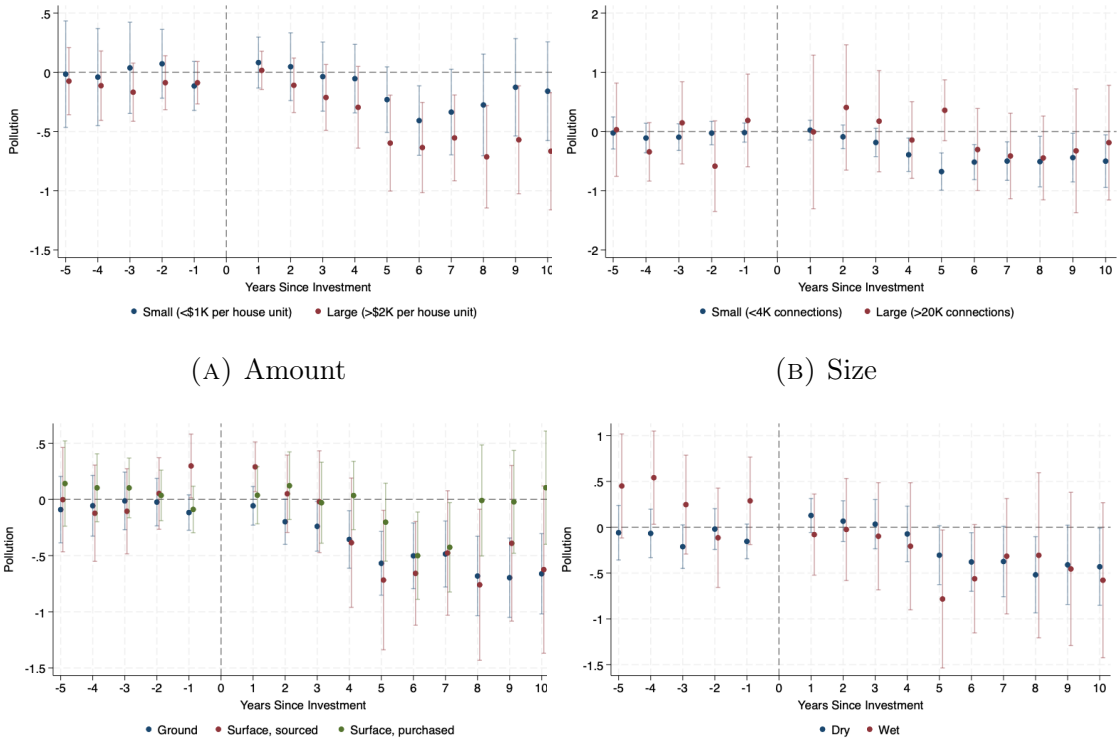
presents the distribution of the log amount of financing through government loans or other long-term debts, based on the 1,074 investment cases where such financing was used. It is notable that only 54% (=1,074/1,979) of realized investment cases involved financing, while the rest is through federal/state grants. These external funding is dominant for smaller investments; the median financing size is $6.45 million with the mean being $38.23 million, both larger than the corresponding statistics for the total investment size.

Panel (B) of Appendix Figure C9 shows that the coupon rate also differs across local governments, depending on the debt type (federal/state government loans vs. other long-term debts like municipal bonds). For federal/state government loans, zero interest rate is not uncommon, with the average rate being 1.32% and the maximum 2.57%. These loans are largely subsidized, as the interest rates in other long-term debts are much higher, with the mean of 3.49% and the maximum of 10.60%.

APPENDIX D. ESTIMATION OF POLITICIANS' PREFERENCES

D.1. **Estimation details.** For a given parameter vector $\boldsymbol{\psi}$, we solve the policy-maker's dynamic program by value function iteration on a discretized $(\Delta r, \Delta q)$ state space representing rate and expected pollution investment stock variables. These stocks are repleted when investment occurs and depreciate over time to match the

FIGURE C8. Investment Effectiveness Across Water System Types



(A) Amount

(B) Size

(C) Water Source

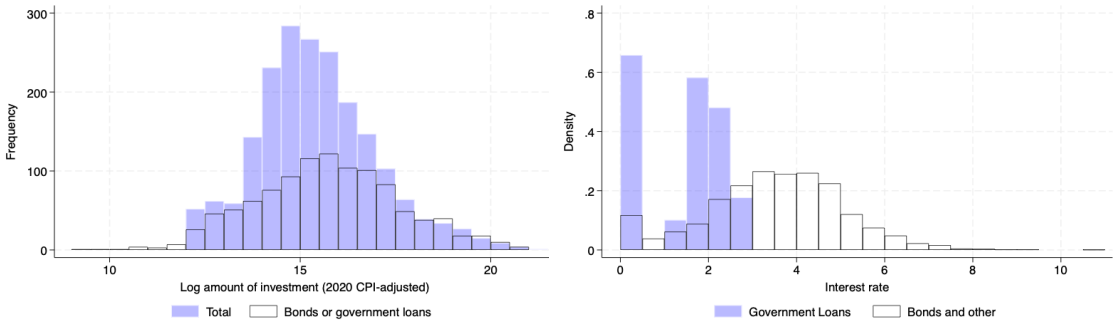
(D) Precipitation

Notes: Each plot presents the coefficient estimates of the dynamic effects of investment on water pollution, using specification (9), for two sub-samples of the data: below-25th percentile and above-75th percentile groups, based on the amount of the investment (Panel A), the size of the system as measured by the number of connections (Panel B), and the average annual amount of rain during the period of study (Panel D), as well as the three groups by the source of water (groundwater, surface water drawn within the jurisdiction, or purchased surface water; Panel C).

shape of our event studies estimated in Figure 3. We assume that the governing body is myopic on other water systems states x_{st}, i.e. that it does not anticipate future changes in voter composition or weather induced pollution or funding generosity.[1] In each state $(\Delta r, \Delta q)$, we first construct residents utility using the estimated

[1]We've also experimented with a perfect foresight assumption that combined an infinite horizon final period with backward induction, and found similar estimates. Both of these approaches provide a parsimonious way to focus on the dynamics of rate and pollution induced by investment while still capturing local differences in demographics and natural amenities. Given the many variables included in x_{st}, incorporating beliefs over all these states would substantially increase the size of the state space.

FIGURE C9. Heterogeneous Investment Costs



(A) Amount (B) Interest Rate

Notes: Panel (A) presents the distribution of the log amount of investment in 2020 CPI-adjusted dollars, based on the observed 1,979 investment cases (in blue) and overlays the distribution of the log amount of financing through government loans or other long-term debts, based on the 1,074 investment cases where such financing was used (in white). Panel (B) shows the distribution of the coupon rates by the debt type (federal/state government loans in blue vs. other long-term debts like municipal bonds in white).

willingness-to-pay $u\big(q_{st}, r_{st}; \omega(y_i)\big)$ and aggregate resident utilities into the policymaker's flow payoff using political weights and turnout:

$$\Pi_{st}(\Delta r, \Delta q; \boldsymbol{\psi}) = \sum_{y} \lambda(\rho_s, y, x_s; \boldsymbol{\psi})\, \tilde{\pi}(y)\, u\big(q_{st}(\Delta q), r_{st}(\Delta r); \omega(y_i)\big),$$

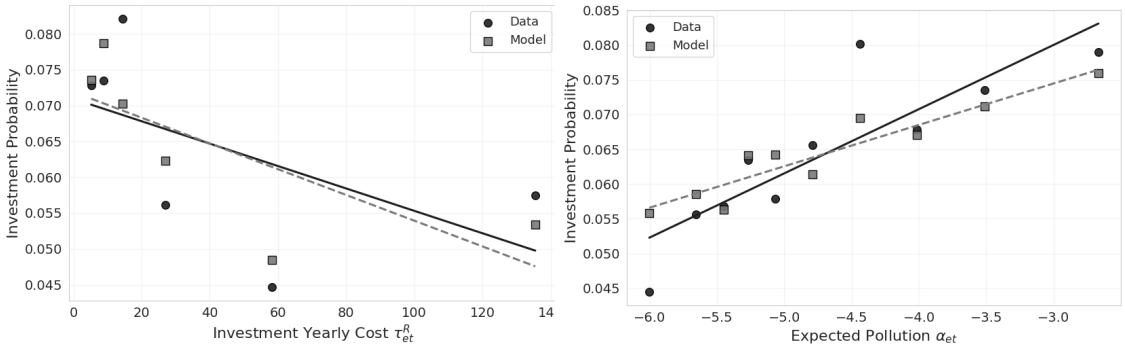
where $q_{st}(\Delta q)$ and $r_{st}(\Delta r)$ denote the mappings from the grid states $(\Delta q, \Delta r)$ into current pollution and rates. Note that critically, we allow the average election contestedness in system s, ρ_s to influence the accountability weight $\lambda(.)$. The law of motion for the discretized state variables implied by our event study estimates are: for $a_{st} \in \{0, 1\}$,

$$\Delta q_{s,t+1} = \rho^q_{st}\, \Delta q_{st} + a_{st}\, \tau^q_{st}, \qquad \Delta r_{s,t+1} = \rho^r_{st}\, \Delta r_{st} + a_{st}\, \tau^r_{st}.$$

Given these primitives and the private cost of investing $\kappa(s_t; \boldsymbol{\psi})$, we solve the optimal policy of each governing body, in each time period using value function iteration. We thus recover the choice-specific value difference $\Delta V_{st}(\Delta r, \Delta q; \boldsymbol{\psi})$ and implied conditional choice probabilities

$$P_{\boldsymbol{\psi}}(a_{st} = 1 \mid \Delta r, \Delta q; x_{st}) = \frac{\exp(\Delta V_{st}(\Delta r, \Delta q; \boldsymbol{\psi}))}{1 + \exp(\Delta V_{st}(\Delta r, \Delta q; \boldsymbol{\psi}))},$$

FIGURE D10. Model Fit on Investment Patterns



(A) Costs (B) Pollution

Notes: This figure shows the model fit along two specific dimensions. Panel (A) shows that the model reproduces the decreasing relationship observed in the data between investment probability and the magnitude of water rate increases from investment. Panel (B) shows that the estimated model successfully reflects that investment is more likely to occur in system-years where expected pollution is likely to be higher, for exogenous reasons such as weather and the water source used.

which we evaluate at observed states and use to construct the log-likelihood

$$\ell(\boldsymbol{\psi}) = \sum_{s,t} \left[a_{st} \log P_{\boldsymbol{\psi}}(a_{st} = 1 \mid \Delta r_{st}, \Delta q_{st}) + (1 - a_{st}) \log \left(1 - P_{\boldsymbol{\psi}}(a_{st} = 1 \mid \Delta r_{st}, \Delta q_{st})\right) \right].$$

We maximize the average log-likelihood using gradient-based optimization, with automatic differentiation and GPU acceleration using PyTorch.