INEFFICIENCIES IN LOCAL INFRASTRUCTURE: EVIDENCE FROM DRINKING WATER IN CALIFORNIA

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Abstract. Local elected officials oversee critical infrastructure investments, yet limited electoral competition may weaken the alignment between local policy and residents' welfare. Using new data on investment and performance for California's drinking-water municipal utilities, we document that jurisdictions with weaker electoral competition exhibit higher water pollution and respond less 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 administrative costs against re-election prospects when choosing investments. Residents' preferences are identified from housing transactions at jurisdiction boundaries and electoral incentives are identified by exploiting variation in voter welfare gains from investment. Counterfactual increases in accountability reallocate investment toward high-need jurisdictions, reduce pollution, and generate progressive welfare gains. Limited accountability also dampens the effectiveness of federal and state subsidies by diverting funds away from high-return investments.

1. Introduction

Infrastructure investment in the United States has become a growing concern as spending struggles to keep pace with an aging capital stock. Local governments oversee roughly half a trillion dollars in annual transportation and water infrastructure spending, yet the accountability of their elected officials is often questioned (Berry, 2009). When electoral oversight is limited, public investment decisions may diverge from residents' needs. This paper quantifies how electoral accountability affects the

Date: March 9, 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.
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allocation of infrastructure investment and the effectiveness of large-scale public subsidies, and evaluates the resulting welfare consequences.

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. Drinking water systems are typically governed by locally elected officials, but these offices are frequently uncontested, and incumbents win most contested races.

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 from 2003 to 2022. 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 cost of time and efforts required to plan and implement the investment. 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 voters' aggregate willingness to pay for water quality improvements. 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 implementation costs.

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 jurisdictions 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 jurisdictions 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 jurisdictions where marginal benefits are lower. These changes translate into economically meaningful welfare gains: in the most pollution-prone jurisdictions, 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 jurisdictions 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 investment decisions by locally elected officials. Second, we directly quantify electoral accountability by estimating how policymakers' responsiveness to constituent

[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).

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.

Finally, this paper builds on a long tradition of using structural models to recover preferences for local amenities, including schools (Bayer et al., 2007), air pollution (Tra, 2010; Bayer et al., 2016), transportation networks (Barwick et al., 2024). We contribute to this literature by introducing a new endogenous channel for the evolution of these local amenities. Rather than arising from residential resorting, as in Bayer et al. (2016); Almagro and Domínguez-Iino (2025), changes in local amenities in our setting are driven by endogenous infrastructure investment decisions made by local policymakers, whose responsiveness to residents' preferences depends on electoral incentives.

2. Drinking water infrastructure

This section describes the institutional environment governing drinking-water infrastructure investment in California. We first outline the basic operations of municipal water utilities. We then discuss how these investments are financed, emphasizing the central role of water service revenues. Next, we describe the local governance structure and electoral institutions under which water utilities operate. Finally, we review the regulatory framework established by the Safe Drinking Water Act and the federal and state subsidy programs that support infrastructure investment. Together, these institutional features shape both the costs and political incentives associated with drinking-water infrastructure provision.

2.1. **Water utility operations.** 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.

2.2. **Financing water infrastructure.** 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 reflects the institutional structure of municipal water operations as enterprise utilities that operate separately from general government finances. Unlike other government services that are supported by tax revenue, water services are funded through user charges, and 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. Importantly, this implies that the relevant opportunity cost of water infrastructure investment arises from the increase in user charges needed to finance the project, rather than from foregone spending on other municipal services such as policing or education.

In California, water rates 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]

2.3. **Local governance and electoral institutions.** 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).

2.4. **Water quality regulation and infrastructure subsidies.** The 1974 Safe Drinking Water Act (as amended in 1996) regulates about 90 pollutants by maximum

[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).

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).

In addition to the regulatory standards for drinking water quality, the federal and state governments has provided substantial financial support to help local systems meet these requirements. Over the past 25 years, the Drinking Water State Revolving Fund (DWSRF) has served as the primary vehicle for this support, providing more than $40 billion in subsidized financing since its inception in 1997 (EPA, 2019). This funding was expanded through the American Recovery and Reinvestment Act of 2009, and more recently, the Infrastructure Investment and Jobs Act of 2021.[3] In California, four statewide general obligation bond measures—Propositions 50 (2002), 84 (2006), 1 (2014), and 68 (2018)—have authorized billions of dollars for water quality and supply investments. More recently, California Senate Bill 200 (2019) created the Safe and Affordable Drinking Water Fund, which provides approximately $130 million annually to support safe drinking water projects.

By channeling large amounts of capital through low-interest loans and grants, these programs reduce the effective cost of drinking-water infrastructure investment, implying that observed water-rate increases capture only a fraction of the true economic cost of providing safe drinking water. Importantly, the availability and intensity of such subsidies vary considerably across jurisdictions and over time, creating meaningful variation in the effective cost of infrastructure investment faced by local governments and their residents. For example, 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]The latter allocated approximately $50 billion for water infrastructure nationwide, including substantial additional funding for the DWSRF, although this expansion falls outside our study period.

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, 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).[4]

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).[5]

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).[6] We identify an investment event when the combined total of debt financing and external funding per year exceeds a minimum threshold.[7] This approach relies on the premise that municipalities rarely fund infrastructure projects entirely from internal resources.[8]

Table 1 summarizes key variables for the cities and special districts separately. Based on our investment measure, Panel A reveals that infrastructure projects are

[4]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.

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

[6]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.

[7]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.

[8]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.

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, remain a secondary source of funding.[9] 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,

[9]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).

mean values demonstrate substantial right-skew, at \$43.7 million and \$16.7 million 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.[10] 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:

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

[10]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.[11] 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 and constraints of local governments 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 collected from county election websites. We further complement these sources with individual-level voter registration data from L2, which provides the universe of registered voters and tracks turnout across federal, state, and local elections. The L2 data also contain demographic and socioeconomic attributes, such as income proxies and age, allowing us to characterize the composition of the electorate and construct measures of electoral participation and voter characteristics at the jurisdiction level.

To measure local governments' administrative capacity, we use the California Government Compensation (GCC) database, which provides comprehensive information on public employees across all local governments, including job titles, compensation, and employment status. These data allow us to construct measures of governmental manpower and, in particular, the presence of highly compensated and skilled employees who may facilitate the planning and implementation of capital investments in drinking water infrastructure.

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.

[11]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.

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 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 j during year t, as defined in (1), varies with the jurisdiction's overall level of electoral competition, e_j. The specification includes jurisdiction and period fixed effects, μ_j and ϕ_t, respectively, as well as time-varying jurisdiction-level demographic characteristics, water system attributes, and precipitation, collected in the vector $\mathbf{x}_{j,t}$.

$$q_{jt} = \beta_e e_j + \beta_x \mathbf{x}_{j,t} + \mu_j + \phi_t + \varepsilon_{jt}, \tag{2}$$

[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 ≤ 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).

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 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.

A natural question is whether drinking-water quality is sufficiently salient to influence local electoral outcomes. Surveys indicate high public concern about drinking-water pollution, and in California advocacy organizations such as the Community Water Center Action Fund and Clean Water Action California seek to raise the political salience of drinking-water issues through voter education, candidate endorsements, and grassroots campaigns. Consistent with this possibility, our results suggest that electoral competition acts as a disciplining force that aligns officials' 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.

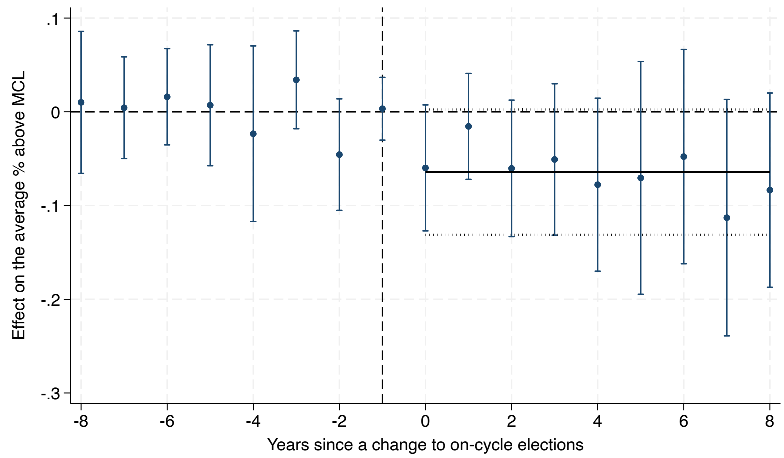
4.1.2. *Event study evidence.* A causal interpretation of the patterns in Table 2 is limited because electoral competition is an equilibrium outcome. For example, poor services can trigger backlash and attract more candidates, increasing competition. To address this concern, we exploit within-jurisdiction 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

[13]Our data confirm this result (Table B1 in Appendix B). 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 (Table B2 in Appendix B), 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.

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_{jt} 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. Overall, moving elections on-cycle is associated with sustained reductions in drinking water pollution.[15]

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 A1 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 measure to state and federal funding. Specifically, we define the expected grant intensity for jurisdiction j in year t, z_{jt}.[16] First, we estimate government-level exposure to federal and state funding, p_j, 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_{jt}, 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 $z_{jt} = p_j g_{jt}$.

Table 3 shows that greater funding availability increases the probability of investment within government. Column (1) indicates that a one-unit increase in z_{jt}

[15]Because infrastructure investments take time to affect water quality, the relatively rapid response likely reflects anticipatory investment decisions. Changes in election timing are often determined prior to the first election under the new schedule, allowing policymakers to adjust investment plans in advance. Consistent with this interpretation, investment effects on water quality in our data typically appear with a lag of about three years (Figure 3(A)), and, for example, the California Voter Participation Rights Act was enacted in 2015 while many jurisdictions implemented on-cycle elections beginning in 2018. Importantly, such anticipatory responses are consistent with the mechanism we study: if policymakers anticipate greater electoral accountability under on-cycle elections, they may adjust investment decisions in advance of the formal transition.

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

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.

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 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, evaluate how electoral incentives shape investment efficiency, and explore the 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 governing body—e.g., a city council or special district board—comprises several members who face reelection on a staggered cycle. The jurisdiction's residents differ in how they value the trade-off between water quality and rates and, therefore, how they value

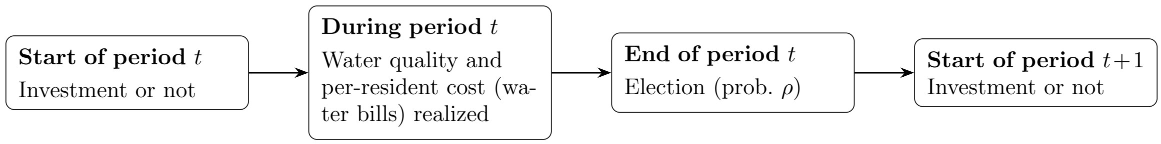


FIGURE 2. Timeline

investment—investment improves water quality, albeit at the cost of increased water rates. Residents therefore decide whether to retain or replace the members standing for election based on current outcomes—water pollution and rates—which reflect the governing body's past investment decisions. These elections are the key disciplining force to align the governing body's investment decision with residents' needs. 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*.[17] Each resident i is characterized by income y_{it}.[18] Let F_{yt} denote the cross-sectional distribution of y in the jurisdiction in period t.

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 per-resident cost of service $r_t \in \mathbb{R}_+$ (covering capital, operations, and maintenance) are publicly observed.[19] At the end of the period, an election occurs with probability ρ. If no election is held, the governing body's composition remains unchanged; otherwise, the next period's membership reflects the vote outcome. If the governing body is approved upon an election (i.e., the members up for reelection are retained), they obtain an extra benefit $\nu \geq 0$. Figure 2 summarizes the timeline.

5.1. **Voting and election outcomes.**

When an election occurs, resident i turns out with probability $\pi_t(y_{it})$; 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.

[17]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.

[18]The framework can accommodate richer, multidimensional resident heterogeneity in principle, but doing so would require additional variation in the data to separately identify preference parameters across multiple resident attributes.

[19]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, we assume that all residents pay the same amount r_t for their water services.

We parameterize resident i's trade-off between water quality and rates by $\omega(y_{it}) \in \mathbb{R}_+$, interpreted as her marginal willingness to pay for quality. The per-period utility from water service is

$$u(q_t, r_t, y_{it}) \;=\; \omega(y_{it})\, q_t \;-\; r_t.$$

Conditional on turning out, resident i 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}:

$$u(q_t, r_t, y_{it}) \;\geq\; v_{it}.$$

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

Let $\tilde{\pi}_t(y_{it}) \equiv \pi_t(y_{it}) / \int \pi_t(y')\, dF_{yt}(y')$ denote turnout weights. The vote share for the incumbent is

$$\tilde{\mu}_t \;=\; \int \tilde{\pi}_t(y)\, F_{v(y)}\big[u(q_t, r_t, y)\big]\, dF_{yt}(y).$$

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

$$\mu(q_t, r_t) \;=\; \frac{1}{2} \;+\; \sigma_\eta\, \sigma_\epsilon \int \tilde{\pi}_t(y)\, u(q_t, r_t, y)\, dF_{yt}(y), \tag{3}$$

provided the arguments lie within the corresponding supports.

5.2. **Water quality and rate dynamics.** Investment in infrastructure presents a fundamental trade-off for the system: it reduces pollution levels but simultaneously increases water rates for residents.

The observed water quality of the system at time t, denoted by q_t, is determined by the interaction between exogenous environmental conditions and the system's infrastructure. The endogenous component of water quality is governed by the investment stock, which encapsulates the cumulative value of past infrastructure improvements. The investment stock l_t^q evolves according to:

$$l_{t+1}^q \;=\; \tau_q a_t + (1 - \delta_q) l_t^q, \tag{4}$$

where the parameter $\delta_q \in [0, 1]$ denotes functional depreciation of the investment stock due to operational wear, and τ_q captures the effectiveness of investment at reducing pollution. Specifically we assume that water quality is a positive continuous variable with conditional mean:

$$\mathbb{E}[q_t] \;=\; \exp(\alpha_t + l_t^q), \tag{5}$$

where α_t captures the expected level of water quality in the absence of investment, reflecting the fundamental quality of the water source which is determined by factors such as local geography and climate conditions.

The per-period, per-resident cost of service reflects both operating expenditures and the recovery of past capital investments. The latter is captured by l_t^r, the per-resident stock of unrecovered investment costs. This state variable represents the accumulated financial obligations or capital charges associated with infrastructure investments. Its evolution depends on parameters τ_r and δ_r, capturing the per-resident cost of new investment a_t and the rate at which these costs are amortized over time.

$$l_{t+1}^r = \tau_r a_t + (1-\delta_r) l_t^r. \tag{6}$$

Under this specification, the present discounted stream of per-resident total cost of investment is $\sum_t^\infty \tau_r (1-\delta_r)^t$. The per-resident water rate then evolves according to

$$r_t = \gamma_t + l_t^r + \zeta_t^r, \tag{7}$$

where γ_t captures the baseline, expected level of water rates, reflecting the jurisdictions' operating costs, in the absence of investment, and the shock ζ_t^r, with $\mathbb{E}(\zeta_t^r) = 0$, represents transitory operations/maintenance costs and is assumed to be serially uncorrelated.

5.3. **The policymaker's investment decision.** The private cost of investment borne by the policymaker is

$$\kappa + \epsilon_t,$$

where ϵ_t is an i.i.d. Logistic shock with unit scale capturing unobserved costs of planning and implementing investment. 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.[20]

The policymaker chooses whether to incur this cost by weighing it against the risk of losing an election, which depends on the residents' welfare gains for investment.

[20]As discussed in Section 2, this cost does not reflect fiscal trade-offs with other municipal spending. Because water infrastructure is financed through water-service revenues rather than the general budget, the policymaker's private cost primarily captures administrative and political effort associated with undertaking the project.

In particular, the policymaker's per-period utility is

$$\rho\mu(q_t, r_t)\nu - (\kappa + \epsilon_t)a. \tag{8}$$

Here the policymaker cares about residents' welfare gains for investment insofar it affects the probability of winning $\mu(q_t, r_t)$. Therefore, these electoral incentives discipline policymakers and help align their actions with constituent welfare. We summarize the institutional features that determine the strength of these incentives– election frequency (ρ), the value of holding office (ν), and the responsiveness of the electorate ($\sigma_\eta, \sigma_\epsilon$)– into a single parameter, λ:

$$\lambda \equiv \nu\, \rho\, \sigma_\eta\, \sigma_\epsilon, \tag{9}$$

so that we can rewrite the policymaker's per-period utility (8) as

$$\lambda \int \tilde{\pi}_t(y) u(q_t, r_t, y)\, dF_{yt}(y) - (\kappa + \epsilon_t)a. \tag{10}$$

We interpret λ as a (reduced-form) measure of political accountability— the weight that the politician places on resident welfare when making investment decisions. A lower value of λ implies a larger institutional wedge between policymaker incentives and constituent welfare, making investment less likely even when residents' aggregate willingness to pay exceeds the policymaker's private cost.

Let β denote the discount factor. Because the policymaker chooses the investment at the start of period t prior to the realization of water service outcomes, the public state is $\mathbf{s}_t \equiv (q_{t-1}, r_{t-1}, l^q_{t-1}, l^r_{t-1})$.

The expected value function can be rewritten as

$$V(\mathbf{s}_t) = \mathbb{E}_{\epsilon_t}\Bigg(\max_{a \in \{0,1\}} \Bigg\{ \lambda \int \tilde{\pi}_t(y)\, \mathbb{E}[u(q_t, r_t, y) \,|\, a, \mathbf{s}_t]\; dF_{yt}(y) \\ - \; a\,(\kappa + \epsilon_t) \; + \; \beta\, \mathbb{E}[V(\mathbf{s}_{t+1}) \,|\, a, \mathbf{s}_t] \Bigg\} \Bigg). \tag{11}$$

Here, $\mathbb{E}_{\epsilon_t}(\cdot)$ averages over the contemporaneous shock ϵ_t; the inner expectations integrate over (q_t, 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 j. In each year t, the policymaker in jurisdiction j chooses an investment decision $a_{jt} \in \{0, 1\}$. Let $\boldsymbol{x}_{jt}$ collect exogenous state variables that affect water quality and costs—e.g., system size and water source—as well as its governance structure (city, dependent special district, or independent special district). Data provide the realized investment decisions a_{jt}, from which the investment stocks (l^q_{jt}, l^r_{jt}) are constructed, as well as water quality q_{jt}, the per-resident water rate r_{jt}, and covariates $\boldsymbol{x}_{jt}$. We also observe the jurisdiction-year-specific distribution of resident income y, denoted by F_{yjt}. In addition, the data include information on voter turnout, elections, government finances related to loans and grants, and housing transactions and property attributes.

6.1. **Overview.** The structural primitives fall into three groups. First, the resident-side primitives are the marginal willingness to pay for water quality $\omega(y)$ and voter turnout rate as a function of their income, $\pi_{jt}(y)$, which we allow to be jurisdiction-year-specific. Second, the government-side primitives are the policymaker's private cost of investing, $\kappa(\boldsymbol{x}_{jt})$ and the weight they place on the constituents' utility $\lambda(\boldsymbol{x}_{jt})$, as well as the intertemporal discount factor, β. We focus on the composite wedge $\lambda(\boldsymbol{x}_{jt})$, which allows us to impose minimal assumptions on the specific mechanism. Finally, the state transition primitives consist of the parameters governing the effectiveness and durability of investment (τ_q, δ_q), the financial cost and amortization of investment (τ_r, δ_r), and the baseline water quality and rate, α_{jt} and γ_{jt}, that we recover for each year and jurisdiction.

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 income-specific turnout $\pi_{jt}(y)$ for each jurisdiction and election cycle from election occurrences and individual-level turnout records. We also recover the state transitions for water quality and rate from data on investment, water quality, and water rate. Using our detailed records on the sources of financing used for capital improvements, we allow investment-induced water rate increases to depend on the expected share of funding coming from state and federal grants rather than from borrowing.

Next, we recover the residents' marginal willingness to pay for water quality, $\omega(y)$. Note that jurisdiction-level investment decisions alone do not disentangle residents' preferences from policymakers' incentives. To address this, we use auxiliary housing transactions data and identify $\omega(y)$ from household sorting. Specifically, we implement a boundary discontinuity design (Black, 1999) nested within an equilibrium housing-demand framework following Bayer et al. (2007).[21]

[21]An alternative approach would be to exploit election outcomes, but these confound the marginal willingness to pay $\omega(y)$ with issue salience and micro-level voting data for local elections are limited.

In the last step, we recover the policymaker's private cost of investing, $\kappa(\boldsymbol{x})$, and the wedges $\lambda(\boldsymbol{x})$ from the observed investment decisions.

6.2. **Water quality transition estimation.** The parameters (τ_q, δ_q) capture the extent to which investment in water infrastructure can improve water quality over time. To recover these parameters, we begin by using a stacked event study approach to assess the impact of an investment across all investment events (Cengiz et al., 2019). For each investment event, e, 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.[22] Since many readings are zero, we use a Poisson regression model with the conditional expectation as follows:

$$\mathbb{E}[q_{jte} | \{I^k_{jte}\}^{16}_{k=-3}, \boldsymbol{x}_{jt}] = \exp\left(\sum_{-3}^{16} \phi^q_k I^k_{jte} + \phi^q_x \boldsymbol{x}_{jt} + \psi^q_{je} + \mu^q_{te} + \eta^q_{jte}\right), \tag{12}$$

where q_{jte} is the water quality measured at the drinking water facility of government j at year t, included in the e^{th} investment event dataset. The treatment dummy I^k_{jte} equals 1 if the investment event e occurred k years from year t by government j, and the specification includes event-specific government fixed effects and event-specific time effects, to ensure that identification comes from within an event.

We then map the estimated event study coefficients to the geometric decay process described in Equation (5) via non-linear least squares (NLS). Specifically, τ_q captures the initial impact of investment on pollution, while δ_q matches the persistence of the effect over time.

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 jurisdiction-year level as a function of jurisdiction 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 expected

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

level of pollution in the absence of investment, α_{jt}:

$$\mathbb{E}[q_{jt}] = \exp\left(\alpha_{jt} + \sum_{e}\sum_{k=0}^{t-t_e} \hat{\phi}_k^q I_{jte}^k\right). \tag{13}$$

To determine the initial investment stock l^q, we exploit the fact that depreciation, governed by δ^q, limits the persistence of past investments. As a result, only a finite number of pre-sample periods must be considered to capture investments that may still affect current pollution and water rates. We thus construct investment histories for the period $1990-2002$ to initialize l_{jt}^q in $t = 2002$ for every jurisdiction. We follow the same approach to initialize the investment stock l_{jt}^r, which is used to estimate the water-rate transition process.

6.3. **Water rate transition estimation.** To estimate how investment costs are 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. To allow for heterogeneity in federal and state funding, which substantially reduces the share of investment costs borne by residents, we estimate the water rate transition process separately for two groups based on the predicted borrowing amount per household. The borrowing among per household is predicted using the relationship between the observed borrowed among in realized investments and the government characteristics considered by federal and state authorities in allocating funds, such as income and population served.

Then, we divide the sample between high- and low- borrowing (respectively above and below median amount per household) and, denoting the two groups by b, for each investment event e we separately estimate

$$\log r_{jte} = \sum_{k=-3}^{16} \phi_k^r I_{jte}^k + \phi_x^r \boldsymbol{x}_{jt} + \psi_{je}^r + \mu_{te}^r + \eta_{jte}^q, \tag{14}$$

where r_{jte} is the per-resident water rate of government j in year t within the e^{th} event dataset, and $\phi_x^r \boldsymbol{x}_{jt}$ captures changes in ongoing maintenance and operating costs. Then, we use the estimated event study coefficients to calibrate a jurisdiction-year-specific parameter τ_j^r. For each borrowing group b, we calculate the average treatment effect on the log water rate over the first five years following an investment event. We then set τ_j^r equal to this average log- increase multiplied by the baseline water rate level.

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).[23] By nesting a boundary discontinuity design (Black, 1999) within an equilibrium housing demand model, we leverage variation in water quality across public water systems.[24]

Consider a potential buyer i with income y_{it} choosing among houses h in a local housing market in year t. Houses vary in annualized transaction prices p_{ht}, physical and neighborhood characteristics $\boldsymbol{x}_{ht}$, the commuting distance to the buyer's workplace, d_{iht}. Each house is also located in a jurisdiction that supplies drinking water, denoted by $j(h)$, which determines the water quality $q_{j(h)t}$ and the annual water rate $r_{j(h)t}$. Houses may additionally differ in an unobserved quality component ξ_{ht}, which may be correlated with price. To account for time-varying unobserved neighborhood attributes that may influence both water pollution and unobserved quality, we include boundary-year fixed effects, $\psi_{b(h)t}$. These fixed effects are defined for houses located within 500 meters of water service jurisdiction boundary $b(h)$ in year t, and capture difference in amenities (other than the difference in water quality and rates) across jurisdictions.[25]

Buyer i's utility from purchasing house h in year t is

$$U_{iht} = \boldsymbol{\theta}_x \boldsymbol{x}_{ht} + \theta_q(y_{it}) q_{j(h)t} - \theta_p(y_{it})(p_{ht} + r_{j(h)t}) + \theta_d d_{iht} + \psi_{b(h)t} + \xi_{ht} + \varepsilon_{iht}, \quad (15)$$

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 sensitivities to water quality and housing costs—the sum of house price and water rate—to vary with buyer income. The idiosyncratic taste shocks ε_{iht} are assumed to follow the Type I extreme value distribution.

[23]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.

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

[25]Ideally, each boundary area lies within a single jurisdiction, allowing the fixed effect to capture factors such as school quality and neighborhood policies (e.g., zoning or crime prevention). Because this condition may not always be met in the data, as a robustness, we additionally control for school fixed effects and city/county government fixed effects, exploiting boundary areas that span multiple school districts or jurisdictions.

Given the demand specification in equation (15), we define resident i's marginal willingness to pay for quality, which varies with income y_{it}, as

$$\omega(y_{it}) = \frac{\theta_q(y_{it})}{\theta_p(y_{it})}.$$

This ratio converts the utility gain from water quality improvements into dollar units, representing resident i's willingness to pay for a marginal increase in water quality.

Estimation of the house demand parameters in equation (15) then proceeds in two stages, following Bayer et al. (2007). In a first step, we estimate the house-specific utilities, denoted by δ_{ht}, along with the non-linear parameters governing buyer heterogeneity, via maximum likelihood. The estimation sample 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 C3. Appendix C provides 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 by regressing the estimated mean utilities δ_{ht} on house characteristics, prices, and water quality. Due to the endogeneity of house prices, we instrument for prices using a variant of BLP instruments (Berry et al., 1995).[26] Note that we allow residents preferences to only depend on the current level of water quality and rates. Therefore, we abstract away from residents dynamic considerations, as our estimates capture flow willingness to pay for current water quality levels. As shown by Bishop and Murphy (2019), if pollution is mean-reverting, as would be expected if severe pollution triggered abatement measures, our estimates would then be biased downward, thus leading us to understate the welfare gains from drinking water quality improvements.

6.5. **Estimation of the policymaker's cost and wedge.** Identification of the flow payoff difference associated with investment follows standard arguments in the dynamic discrete choice literature (Rust, 1987; Hotz and Miller, 1993).[27]

[26]Specifically, we capture the influence of nearby houses while excluding those in close proximity, exploiting variation in the exogenous characteristics of alternatives without picking up unobserved local amenities that are shared within small geographic areas. We provide additional details in Appendix C.

[27]Note that 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). Moreover, we assume that the policymaker has rational expectations about the endogenous state variables (l^q, l^r) , where beliefs about future states are identified from observed transitions in water quality and rates, as discussed above. For tractability, although this could be relaxed, we assume that the policymakers do not anticipate future changes in exogenous state variables.

We then separately identify the payoff parameters $\lambda(\boldsymbol{x}_{jt})$ and $\kappa(\boldsymbol{x}_{jt})$ using the following sources of variation in the data. The wedge $\lambda(\boldsymbol{x}_{jt})$ is identified from variation in residents' aggregate willingness to pay that shifts the political payoff from investment while leaving the policymaker's private implementation cost unchanged. Specifically, the responsiveness of investment to these changes in residents' willingness to pay for investment pins down $\lambda(\boldsymbol{x}_{jt})$. In contrast, $\kappa(\boldsymbol{x}_{jt})$ is identified from residual variation in investment decisions after accounting for these welfare shifts and the responsiveness of policymakers captured by $\lambda(\boldsymbol{x}_{jt})$.

To generate variation in electorates' aggregate willingness to pay for investment—the component of resident welfare that enters the policymaker's political payoff—we exploit changes in voter turnout, the income composition of the electorate, and residents' valuation of investment, which depends on the underlying water quality and rate structure (α_{jt} and γ_{jt}). These determinants evolve due to institutional changes such as shifts from off-cycle to on-cycle elections, demographic changes such as migration, and fluctuations in environmental conditions affecting pollution levels. Policymakers observe these state variables and anticipate their implications for residents' welfare when making investment decisions.

With these components in place, the policymaker preferences are then estimated by maximum likelihood, solving the dynamic program at each parameter guess with a nested fixed-point algorithm (Rust, 1987).

In estimation, we treat $\kappa(\boldsymbol{x}_{jt})$ and $\lambda(\boldsymbol{x}_{jt})$ as functions of persistent jurisdiction-specific institutional features that do not respond to short-run demographic or pollution fluctuations. In particular, we specify policymakers' private investment cost, $\kappa(\boldsymbol{x}_{jt})$, which reflects the jurisdiction's *capacity* to plan implement capital projects, as a linear function of (i) the log number of employees engaged in drinking-water-related activities and (ii) an indicator equal to one if the system serves a small population (fewer than 10,000 residents). To discipline the welfare weight $\lambda(\boldsymbol{x}_{jt})$, we construct an index of electoral accountability for each jurisdiction, denoted by A_j, and assume

$$\lambda_j = \exp(\lambda_0 + \lambda_1 A_j)\,.$$

Our baseline accountability measure is derived from the frequency of contested elections–the key factor disciplining the policymaker's actions.[28]. Namely, we predict the probability of a contested election based on a Probit model with jurisdiction and year

[28]As a robustness check, we also construct a composite measure that incorporates both the frequency of contested elections and incumbents' vote margins.

Table 4. Heterogeneous Effects of Investment on Water Rate

	Low Borrowing	High Borrowing
Investment Effect	0.016	0.093
	(0.026)	(0.038)

Notes: This table reports the average post-treatment effects on water rates during the first five years following a drinking-water infrastructure investment, separately for jurisdictions with high and low levels of investment size, net of federal or state subsidy support. Standard errors are clustered at the jurisdiction level. The sample covers 2003–2022.

fixed effects and an interaction between an even-year indicator and an indicator for jurisdictions with even-year elections. Our measure of electoral accountability, A_j is the average predicted probability of a contested election over two-year intervals for each jurisdiction j.[29] With this measure of electoral accountability we aim to capture exogenous jurisdictional characteristics arising from entrenched, slow-moving institutional features—such as election cycles and government structure—and therefore not subject to rapid changes.

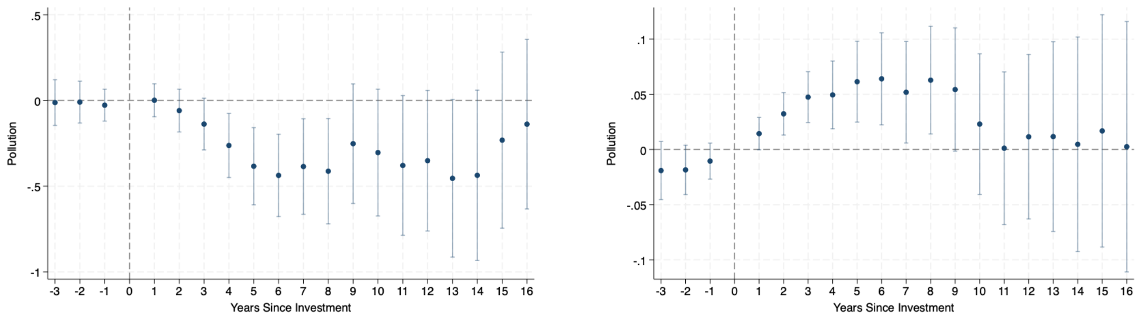
7. Estimation results

This section reports the model estimates. We begin by the estimated water quality and cost dynamics, highlighting the effectiveness capital investments in drinking water infrastructure and the heterogeneous costs borne by residents. We then turn to residents' preferences, summarized by their marginal willingness to pay for quality, $\omega(y)$. Together, these estimates allow us to assess distortions in investment decisions without relying on the policymaker parameters. We then present estimates of the policymaker's capacity cost in planning and undertaking an investment and their responsiveness to resident welfare.

7.1. **Effectiveness and cost of investments.** Panel (A) of Figure 3 plots the event-study coefficients from (12). Relative to the pre-investment baseline, water-system upgrades produce a sharp and statistically significant decline in contamination: in

[29] In the model, λ_{jt} is in principle allowed to depend on a vector of jurisdictional characteristics, $\boldsymbol{x}_{jt}$. In practice, however, identifying a flexible mapping from $\boldsymbol{x_{jt}}$ to λ_{jt} is challenging because it requires sufficient independent variation in the welfare gains from investment across jurisdictions and over time. Much like estimating heterogeneous slope coefficients in a reduced-form regression, a richly parameterized specification would demand substantially more variation than is available in the data. We therefore summarize relevant institutional heterogeneity using the index A_j and parameterize λ_j as a function of this scalar measure.

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.

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 D shows that the effect is fairly uniform across systems with different observable characteristics. Notably, the improvement erodes over time and has vanished in sixteen years. This maps to $\delta^q = 0.878$ and $\tau^q = 1.364$.

Table 4 presents the event-study estimates from (14). 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.

Reflecting the substantial heterogeneity in financing terms (Appendix Figure D8), our model allows the rate impact of investment to depend on the predicted investment amount per household, *net of state and federal grants*. The results in the table show that rate increases are driven almost entirely by the jurisdictions that rely heavily on borrowing to finance drinking-water infrastructure instead of federal and state subsidies ("High Borrowing"). Jurisdictions below the median borrowing level experience virtually no change in water rates, whereas those above the median see an average increase of about 9%.

Table 5. Residents Value Drinking Water Quality Heterogeneously

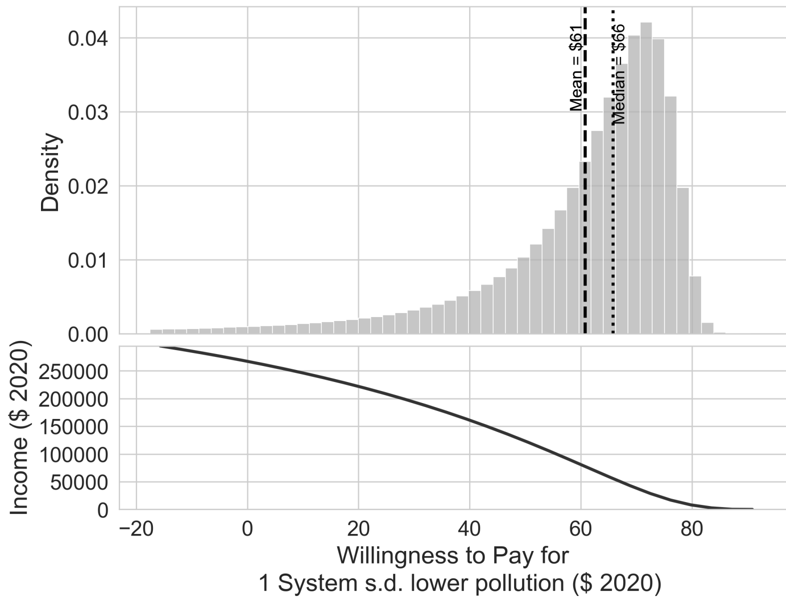
	Without Neigh. Pref. (1)	With Neigh. Pref. (2)
First Stage		
Income (log) × Water pollution	0.008 (0.550)	0.025 (0.115)
Income (log) × House price	36.241 (0.044)	32.116 (0.118)
Distance to workplace	-6.792 (0.040)	-6.799 (0.273)
Income (log) × Block-group income		0.480 (0.236)
Second Stage (IV) - Dep. 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 demand estimation results from the two-stage procedure of Bayer et al. (2007). Standard errors are in parentheses. The specification in column (2) allows preferences for neighborhood income composition to vary with household income by including an interaction between household income and block-group income.

7.2. **Resident preferences for drinking water quality.** Table 5 presents the demand estimates specified in (15), obtained using the two-stage procedure of Bayer et al. (2007). We estimate two specifications, with and without allowing preferences for neighborhood income composition that vary with household income. The results are similar across specifications. As expected, the coefficient on commuting distance θ_d and the coefficient on the composite annualized housing cost (including the water rates) θ_{p0} are both negative: residents prefer shorter commutes and lower housing costs. 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 estimates 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 literature (Bayer

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



Notes: This graph presents the distribution of the willingness to pay $\omega(y)$ for a one standard deviation increase in water quality across households and water systems per year, where the standard deviation is computed across jurisdiction-year, in CPI-adjusted 2020 dollars, using the demand estimates from Column (1) of Table 5. Each system in this graph is represented with an equal population.

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 lower marginal willingness to pay for water quality than lower-income residents.

In Appendix C.2, we provide additional results that confirm that house prices are responsive to water pollution. First, we show hedonic regressions that show 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 C.2 we also present an event study of house prices after an health-based drinking water violation. We find that house prices within a jurisdiction's boundary fall significantly in the three years following the violation, suggesting that water quality is an important amenity valued by residents.

Based on our estimates from Column (1) of Table 5, we provide the distribution of the corresponding annualized willingness to pay for a one standard deviation decrease in water pollution across households, jurisdictions, 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 heterogeneous price and pollution sensitivities, once we account for heterogeneous sorting based on preferences for richer neighbors, we find that willingness to pay decreases with household income.

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:

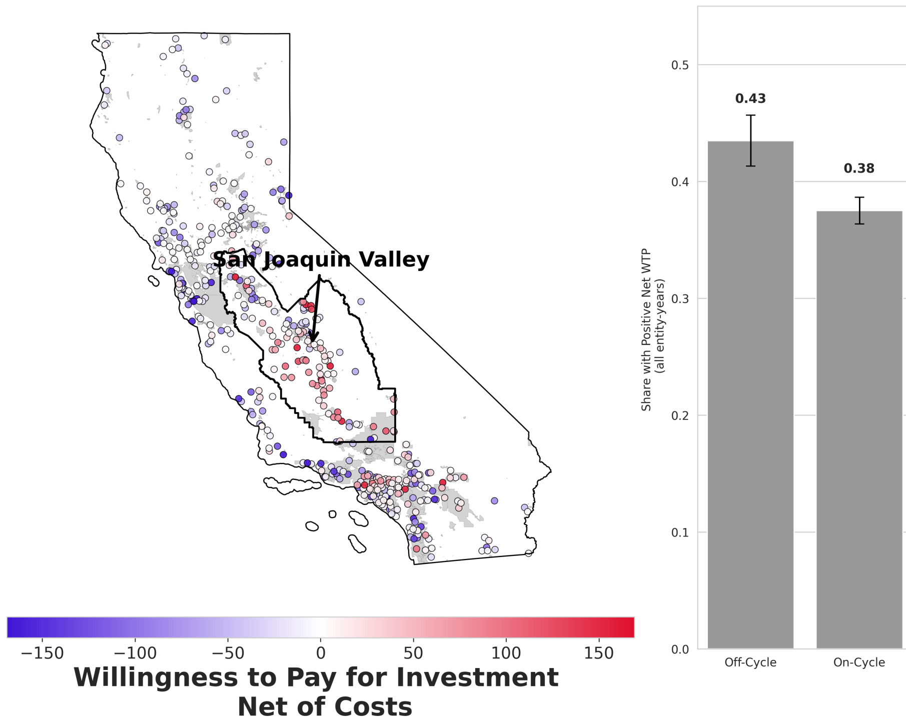
$$\Delta u(q_t, r_t; y_{it}) = \omega(y_{it})\Delta q_t - \Delta r_t, \tag{16}$$

where the changes Δr_t and Δq_t are quantified from the estimates in Section 7.1, and $\omega(y_{it})$ 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.

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.

Figure 5. Net Benefits of Investment across Space and Election Timing



Notes: On the left panel, the figure displays the net investment benefits in 2020 USD as defined in equation (16) 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")

7.4. **How local political factors shape infrastructure investment.** Table 6 presents the estimated parameters governing political constraints in the model. Consistent with the evidence in Section 4, our estimates reveal that when electoral accountability is higher, local officials are more responsive to their constituents welfare. As a result, we show in Appendix Figure E9 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

Table 6. Estimates of the Role of Political Factors in Investment

	Coefficient	Std. Error
κ: Capacity		
Constant	3.276	(0.371)
Number of skilled employees (log)	-0.157	(0.028)
Small-sized	0.454	(0.039)
λ: Weight for constituent welfare		
Constant	-15.874	(3.160)
Electoral accountability index	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 κ is linear in the jurisdiction attributes, and the specification for λ is log-linear in a constant and an accountability measure based on the jurisdiction-specific frequency of contested elections during the period of study.

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).[30]

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 our accountability index is set equal to its highest level in the sample.[31,32] We first simulate the resulting investment paths and their implications for water pollution

[30]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.

[31]Given our definition of the index, this corresponds to the case where governments face consistently contested elections throughout the sample period.

[32]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 (Table B2 in the Appendix B). This episode provides a concrete policy precedent for changes that enhance local electoral accountability.

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 high-accountability 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.[33] As we describe below, overall, our results showcase that greater electoral accountability increases investment in the systems that need it most.

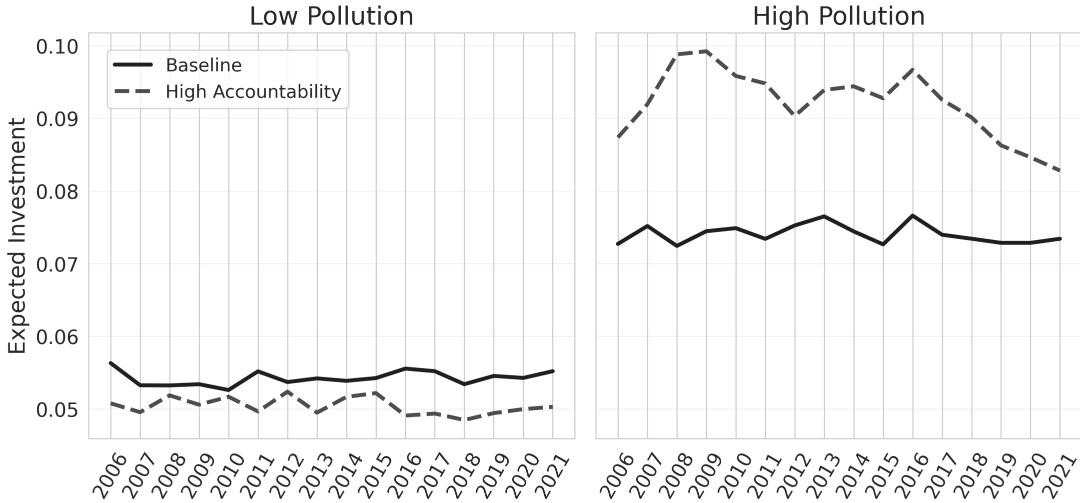
Panel (A) of Figure 6 plots average investment probability under both the observed baseline and the high-accountability counterfactual, split by quintiles of underlying system pollution. We define underlying pollution as $\frac{1}{T}\sum_t \alpha_{jt}$ from the estimated pollution law of motion. Under the high-accountability counterfactual, investment is reallocated toward jurisdictions with the greatest need and modestly reduced in cleaner jurisdictions relative to the baseline.[34] In the most pollution-prone systems (right panel), 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

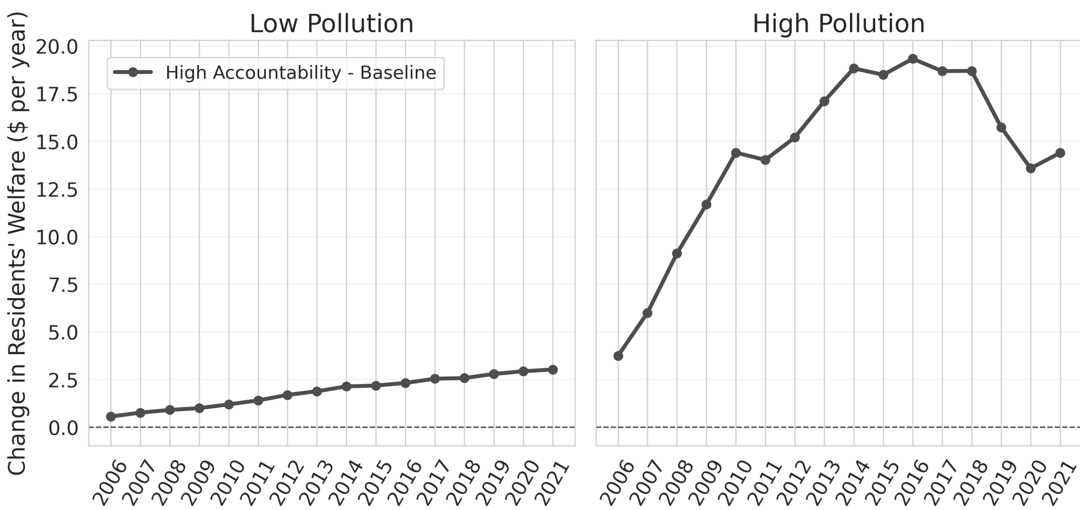
[33]Specifically, we initialize every system to their observed state in 2003 and simulate 1,000 paths of investment, pollution, and water rates until the end of our estimation period, 2022.

[34]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 the accountability index A_j is set equal to its highest level for all jurisdictions. The investment probability is averaged across simulated paths and across governments grouped by quintiles of average expected pollution. For each government j, average expected pollution is defined as the sample mean of α_{jt} in (13). 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 ("High Accountability") 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.

Section 6, our preference parameter estimates may be conservative with respect to the health and economic costs of drinking water contamination.[35]

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 increasing accountability 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 jurisdictions also suffer from weak electoral competition at the local level.[36] As such, even when funding is available, the incentives for local officials in those water boards to address pressing pollution issues may not be strong.[37] 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 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 examine how limited political accountability affects the cost-effectiveness of federal and state funding programs for drinking water infrastructure, possibly undermining the targeting of these programs. Despite a range of federal and state financing initiatives (Section 2.4), funds may not necessarily flow to the jurisdictions that need them most (Hansen et al., 2021). First, we compute the counterfactual water rate increases that would be needed to fund infrastructure investment in the absence of external subsidies.[38] Then, using our estimated model, we solve for the revised optimal investment decisions of local officials across the state space, under both the high-accountability environment, and the observed one.

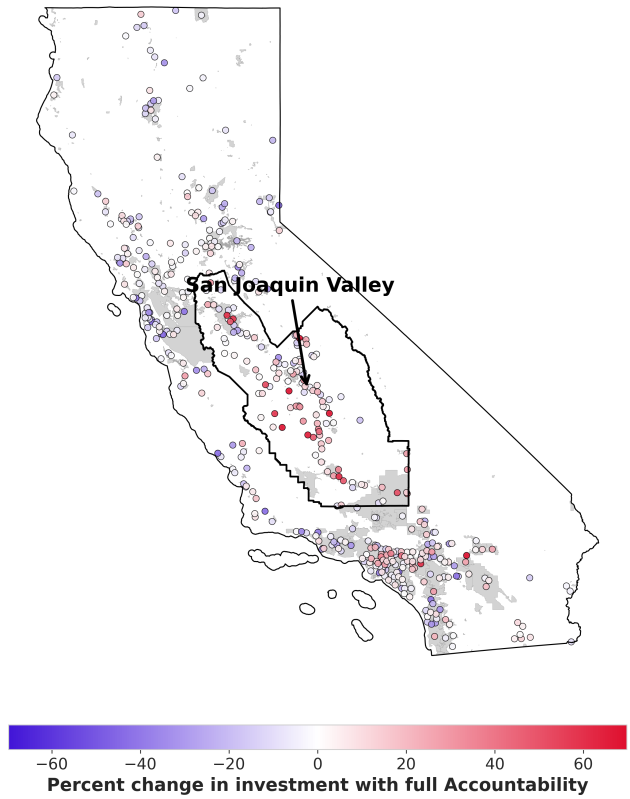
[35]For example, Keiser et al. (2023) report that DWSRF loans and grants reduce mortality among older adults, and Hill and Ma (2022) documents adverse impacts of fracking-induced water pollution on infant health.

[36]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."

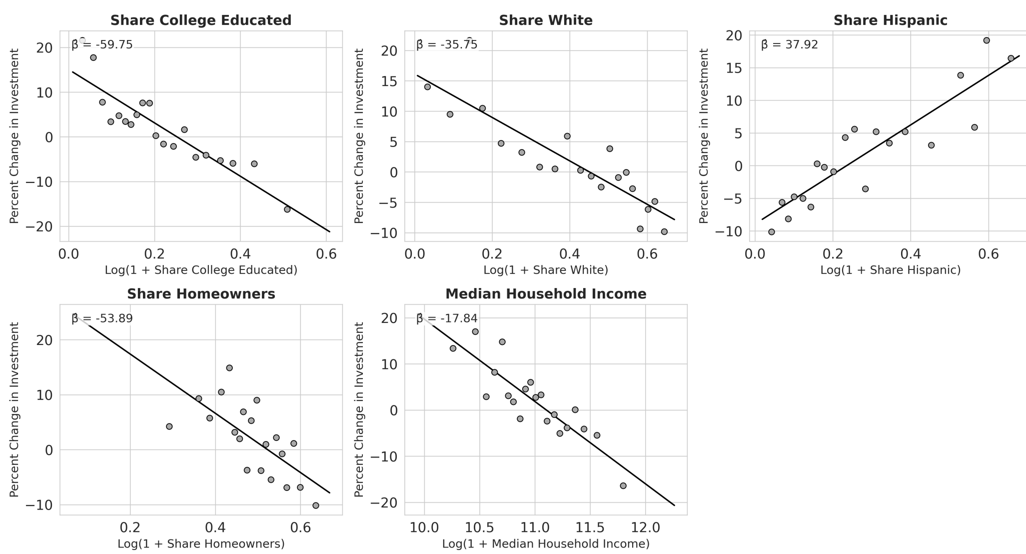
[37]Anecdotally, East Orosi CSD in the San Joaquin Valley accumulated 99 water quality violations over our sample period without undertaking any investment or holding a contested election; the system was placed under a state administrative order in 2022.

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

FIGURE 7. Heterogeneous Effects of Accountability on Investment

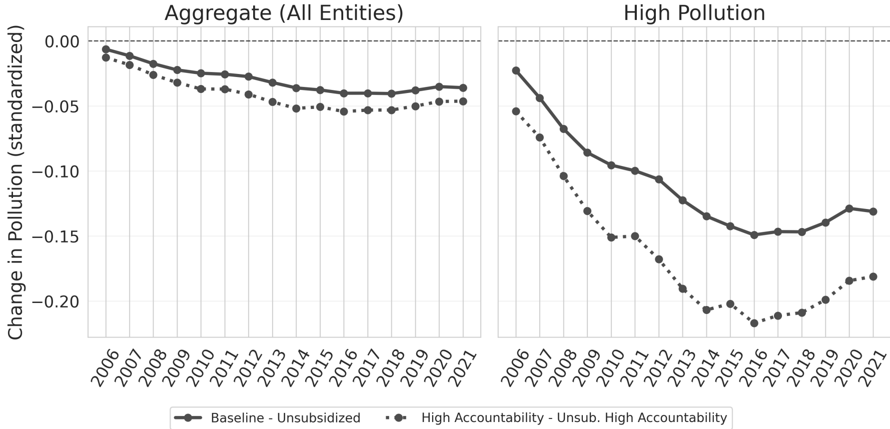


(A) Across Space

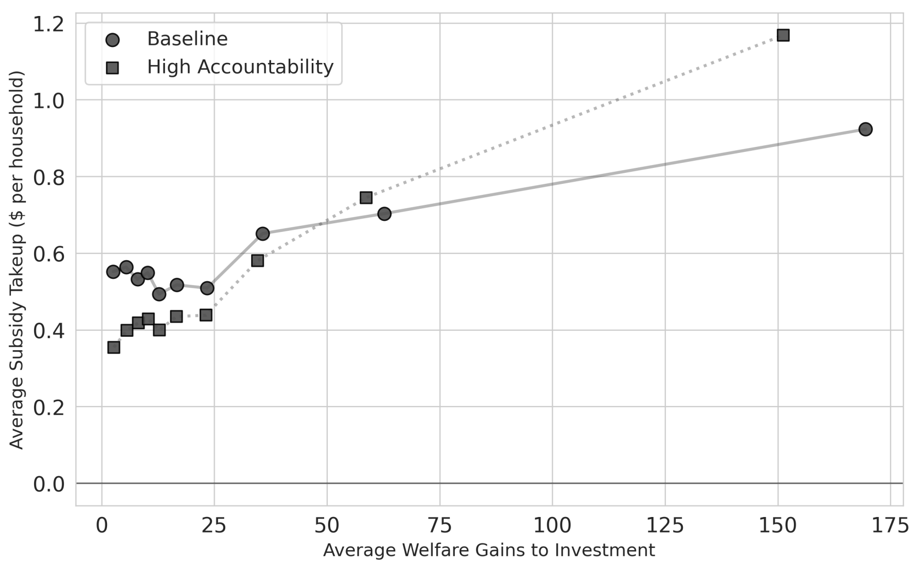


(B) Across Demographics

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 high-accountability scenarios, while holding the subsidy environment constant. "Baseline - Unsubsidized" (solid line) indicates that subsidy effectiveness is simulated at observed levels of electoral accountability. "High Accountability - Unsub. High Accountability" (dashed line) indicates a simulation under a scenario where every election has at least two contenders, so that our accountability index A_j is set to be equal to its highest level in the sample. 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.

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 high-accountability 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 high-accountability 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 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.

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Appendix A. Data Construction

A.1. **Temporal variation in Federal and State funding.** Figure A1 illustrates temporal variation in funding availability for drinking water infrastructure during our study period. The solid line plots the mean of grant amount per capita received by local governments in our sample, constructed from our administrative data combining California's Financial Transactions Reports (FTR), California Debt and Investment Advisory Commission (CDIAC) records, and EPA State Revolving Fund records acquired through FOIA requests. The dashed line summarizes major federal and state funding initiatives identified through our institutional review, including voter-approved bond measures and federal statutory funding expansions. The vertical lines mark the enactment years of these policies. While the two series are constructed from different sources, they exhibit similar time variation, with increases in grant funding following major legislative initiatives. These patterns highlight substantial temporal fluctuations in funding opportunities faced by local governments, which we exploit to identify the responsiveness of infrastructure investment to exogenous changes in funding availability.

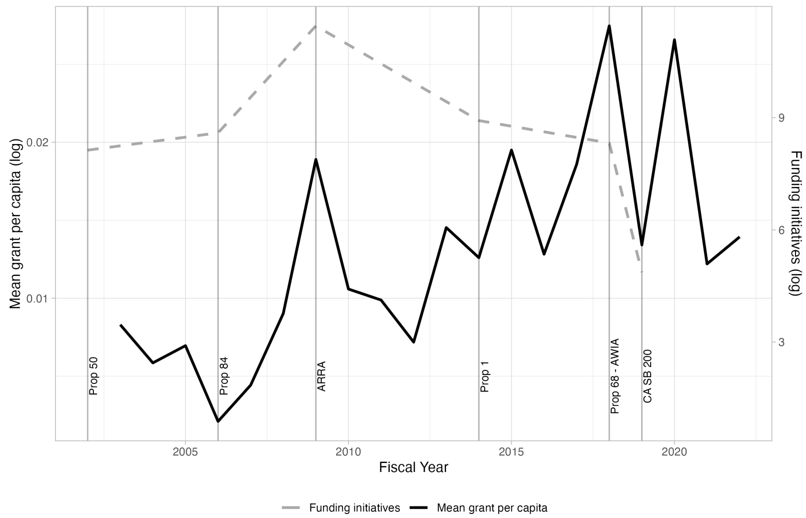
A.2. **Expected grant intensity.** Figure A1 illustrates the evolution of statutory funding initiatives for drinking water infrastructure in California. Over the sample period, several federal and state policies expanded the availability of grant funding, including the American Recovery and Reinvestment Act of 2009 and multiple state bond measures. Consistent with these policy changes, the amount of grant-funded investment per capita in our data increases over time as successive initiatives accumulate. Although the precise eligibility criteria and allocation rules for these programs are not fully observed, the timing of these legislative changes generates substantial temporal variation in funding availability.

To exploit this variation, we construct a measure of expected grant intensity that combines cross-sectional exposure to funding with time-series funding shocks. For each jurisdiction j in year t, we define an indicator for whether the jurisdiction received grant funding,

$$h_{jt} = \mathbb{1}\{\text{Grant}_{jt} > 0\}.$$

While detailed allocation rules are not available, funding programs typically prioritize jurisdictions based on characteristics such as system size, income levels, and financial capacity. To capture systematic differences in exposure to these programs, we estimate a predicted probability of receiving funding, $\hat{h}_{jt} = f(\boldsymbol{x}_{jt})$, using a random forest

Figure A1. 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.

model. The covariates $\boldsymbol{x}_{jt}$ include median household income; an indicator for income below program priority thresholds; an indicator for poor financial health; housing density; an indicator for small water systems; water source type; county; and year. We then define each jurisdiction's average exposure to funding as

$$p_j = \frac{1}{T} \sum_t \hat{h}_{jt}.$$

Next, we construct funding shocks based on realized grant allocations among demographically similar jurisdictions. Specifically, we form a donor pool $\mathcal{P}_j$ using k-means clustering on baseline (2003) characteristics–log population, log income, and log housing density. The median donor pool contains 33 jurisdictions. For each jurisdiction j, the funding shock in year t is defined as

$$g_{jt} = \exp\left(\frac{\sum_{k \in \mathcal{P}_j, k \neq j} \log(1 + \text{Grant Amt}_{kt})}{\sum_{k \in \mathcal{P}_j, k \neq j} 1} \right) - 1,$$

which captures contemporaneous funding variation among comparable jurisdictions while excluding jurisdiction j itself.

Finally, we combine exposure and funding shocks to construct the expected grant intensity measure used in the main text:

$$z_{jt} = p_j g_{jt}.$$

This shift-share measure isolates variation in funding availability arising from aggregate policy-driven funding shocks while scaling it by each jurisdiction's predicted exposure to these programs.

Appendix B. Effects of Election Timing

Appendix Tables B1 and B2 provide additional evidence on how election timing affects electoral competition in our data. Table B1 examines whether elections are contested. Using jurisdiction × two-year observations with jurisdiction and year fixed effects, we find that elections held on-cycle—that is, aligned with statewide or federal contests—are substantially more likely to attract multiple candidates. Across specifications, moving to an on-cycle schedule increases the probability that an election is contested by roughly 12 to 25 percentage points. These patterns are robust to the inclusion of demographic controls and also hold in a subsample focusing on jurisdictions affected by the implementation of the California Voter Participation Rights Act (SB 415, 2015), which required some jurisdictions to move their elections on-cycle beginning in 2018.

Table B2 examines electoral competitiveness using incumbent vote margins. The dependent variable is defined as the minimum of zero and the incumbent vote margins averaged across elections within a jurisdiction × two-year period. We report results for both unbalanced and balanced samples. In the balanced specification, we impute an incumbent vote margin of one in uncontested elections when election information is available from the CEDA database, while the unbalanced specification restricts attention to observed margins. Consistent with the results on contestation, the estimates indicate that on-cycle elections are associated with smaller incumbent vote margins, implying more competitive races. The magnitude of the effect is larger in the balanced panel, reflecting the incorporation of uncontested elections into the measure of electoral competitiveness.

Taken together, these results suggest that aligning local elections with higher-turnout statewide or federal contests is associated with greater electoral competition in our setting, both in terms of the likelihood that races are contested and the margins by which incumbents win.

Table B1. Election Timing and Contested Elections

	Dependent var: Any contested elections			
	Full sample		Subsample	
	(1)	(2)	(3)	(4)
On-cycle election schedule	0.125***	0.117***	0.247***	0.235***
	(0.0326)	(0.0330)	(0.0474)	(0.0506)
Demographic attributes	No	Yes	No	Yes
Year FE	Yes	Yes	Yes	Yes
Jurisdiction FE	Yes	Yes	Yes	Yes
Observations	7,039	6,926	3,230	3,179
R^2	0.562	0.565	0.604	0.606

Notes. This table reports regression results where the dependent variable is an indicator for whether an election is contested. The unit of observation is jurisdiction × two-year period, reflecting that elections typically occur every two years. Standard errors clustered at the jurisdiction level are reported in parentheses. Statistical significance is indicated as follows: $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$. Columns (1)–(2) use the full sample, while Columns (3)–(4) restrict the data to the post-2014 period and excludes jurisdictions that already had on-cycle election schedules prior to 2018, thereby exploiting variation from the implementation of the California Voter Participation Rights Act (SB 415, 2015). Controls include the logarithm of median household income, population, and population density; the shares of White and Hispanic residents and homeowners; and measures of precipitation levels and intensity.

Appendix C. Demand Estimation

C.1. **Estimation approach.**

C.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 C3 below shows the resulting markets.

TABLE B2. Election Timing and Vote Margins

	Dependent var: Incumbents' vote margins			
	Unbalanced		Balanced	
	Full	Subsample	Full	Subsample
	(1)	(2)	(3)	(4)
On-cycle election schedule	0.00487	-0.0739	-0.0955*	-0.152**
	(0.0295)	(0.0573)	(0.0502)	(0.0687)
Demographic attributes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes
Jurisdiction FE	Yes	Yes	Yes	Yes
Observations	2,845	1,155	3,671	1,605
R^2	0.555	0.674	0.474	0.611

Notes. The dependent variable is $\min\{0, \text{incumbent vote margin, averaged across elections}\}$. The unit of observation is a jurisdiction × two-year period, reflecting that elections typically occur every two years. Standard errors clustered at the jurisdiction level are reported in parentheses. $^*p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$. Columns differ by sample definition. "Full" uses the entire sample. "Subsample" restricts the data to the post-2014 period and excludes jurisdictions that already had on-cycle election schedules prior to 2018, thereby exploiting variation from the implementation of the California Voter Participation Rights Act (SB 415, 2015). "Balanced" imputes an incumbent vote margin of 1 in uncontested elections when election information is available from the CEDA database, while "Unbalanced" does not perform this imputation. Controls include the logarithm of median household income, population, and population density; the shares of White and Hispanic residents and homeowners; and measures of precipitation levels and intensity.

FIGURE C2. Geographic Markets used in Demand

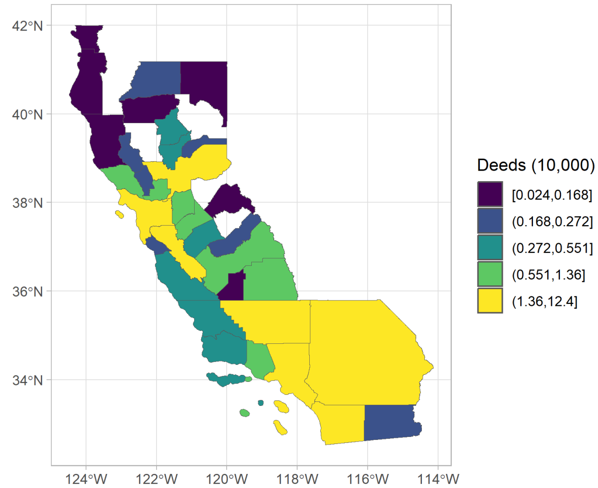
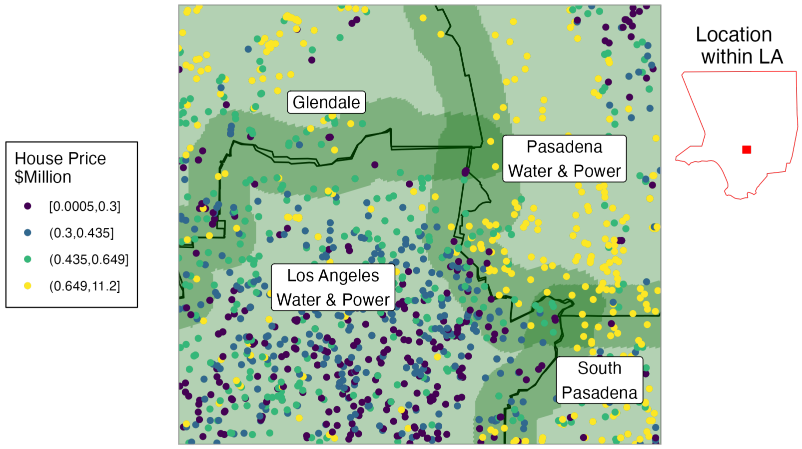


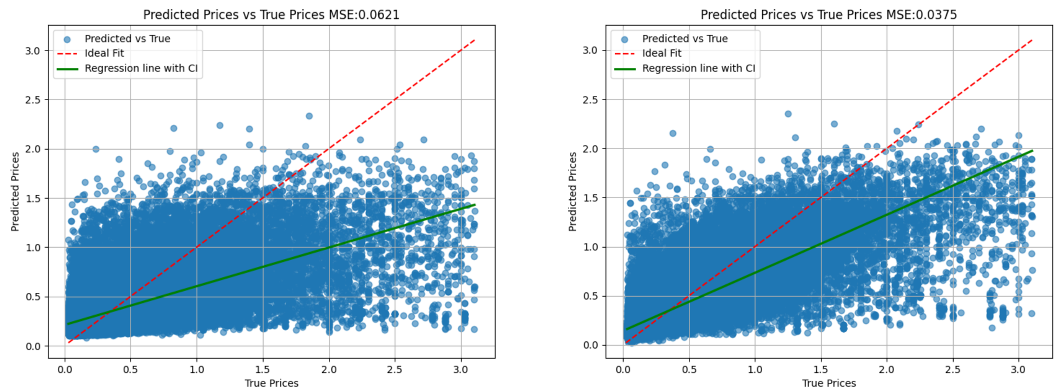
Figure C3. Boundary Discontinuity Design



C.1.2. *Computational details.* As shown in Figure C3, 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 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.

C.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

FIGURE C4. Best Predictor of House Price



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

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 C4 that this instrument 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.

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

$$\log(p_{ht} + r_{ht}) = \beta q_{ht} + \gamma \boldsymbol{x}_{ht} + \psi_{b(h)t} + \rho_s + \mu_g + \epsilon_{ht}, \tag{17}$$

where q_{ht} corresponds to different measures of water quality and where we control for house characteristics and neighborhood demographics, represented by $\boldsymbol{x}_{ht}$. To account for time-varying unobserved neighborhood attributes that influence house prices, we include boundary-year fixed effects, $\psi_{b(h)t}$, which equals one if house h is located within 500m distance of a PWS service area boundary b in year t, $\psi_{b(h)t}$. Ideally, each boundary area lies within a single jurisdiction, allowing $\psi_{b(h)t}$ 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

Table C3. 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

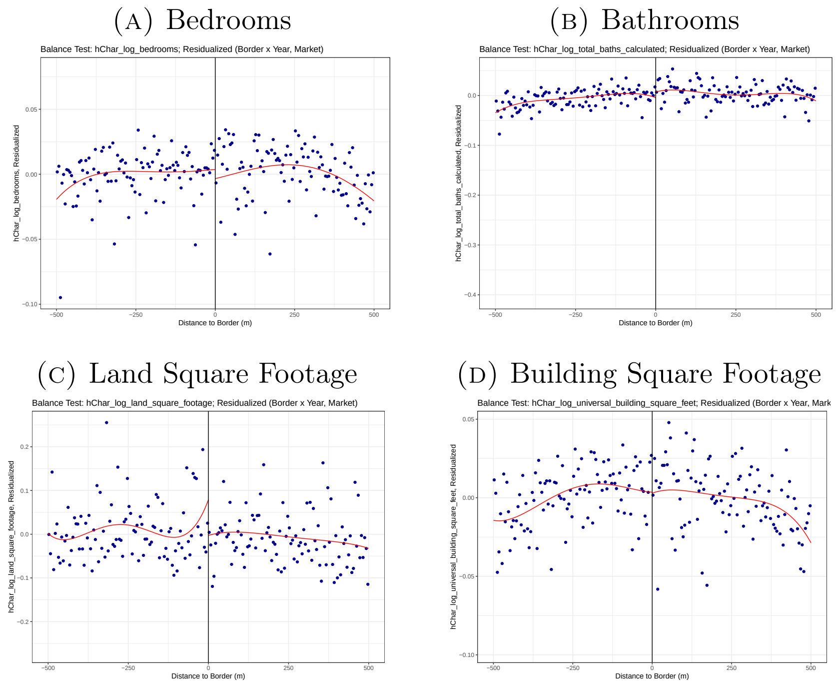
Notes. This table presents OLS estimates based on house transactions in California from 2003 to 2019. Standard errors, clustered at the boundary-year level, are shown in parentheses. Statistical significance is indicated as follows: $^*p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$. The dependent variable 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.

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 C3 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 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'

FIGURE C5. 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.

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 C5 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.

C.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 j with violation event e in year t, and a house by h, we estimate

$$\log p_{hjet} = \beta x_{ht} + \phi_{je} + \mu_t \tag{18}$$

$$+ \sum_{\tau} \delta_{\tau}^{I} \; I_{j,t,e}^{\tau} \mathbb{1}\{e \text{ is Tier I}\} \tag{19}$$

$$+ \sum_{\tau} \delta_{\tau}^{II} \; I_{j,t,e}^{\tau} \mathbb{1}\{e \text{ is Tier II}\} + \epsilon_{hjet} \tag{20}$$

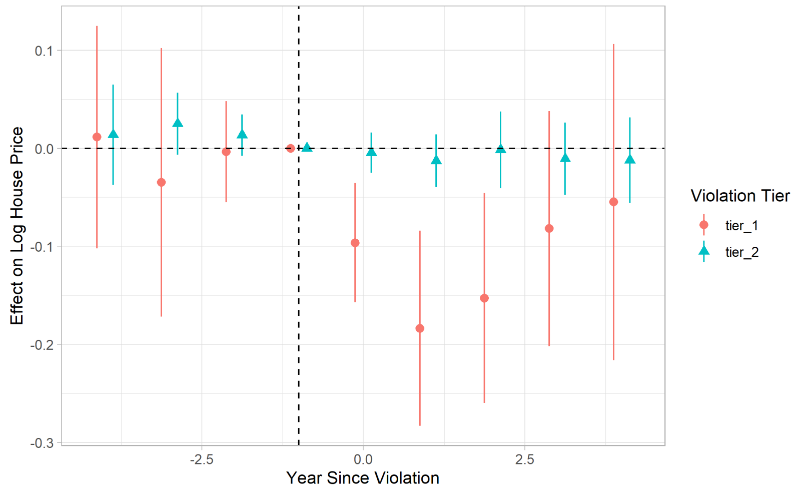
The treatment dummy $I_{j,t,e}^{\tau}$ equals 1 if the violation e occurred τ years from year t by jurisdiction j, and the specification includes event-specific water system fixed effects and year fixed effects, to account for systems with multiple violations over our sample. We control for a rich set of house characteristics x_{hjet}, 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 C6. Consistent with Table C3 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 D. Infrastructure investments: Heterogeneity in effects

D.1. **Homogeneous effects of investment on water quality.** Figure D7 presents the dynamic effects of investment on water pollution, using specification (12), 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).

D.2. **Heterogeneous costs of investment.** Figure D8 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

Figure C6. Event Study: Effects of Violations on House Prices



Notes: This figure presents estimates from the event-study specification 20estimated 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.

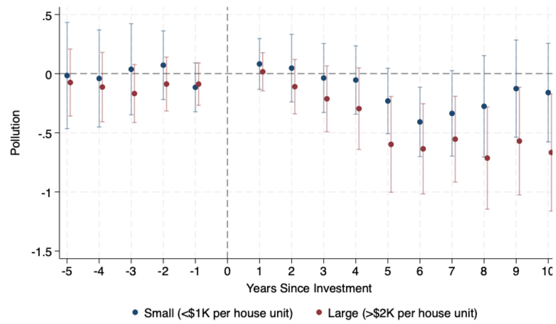
1,979 investment cases. The median investment size is \$4.38 million while the average is \$23.64 million. The panel also presents the distribution of the log amount of financing through government loans or other long-term debts, based on the 1,074 investment cases where such financing was used. It is 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 D8 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%.

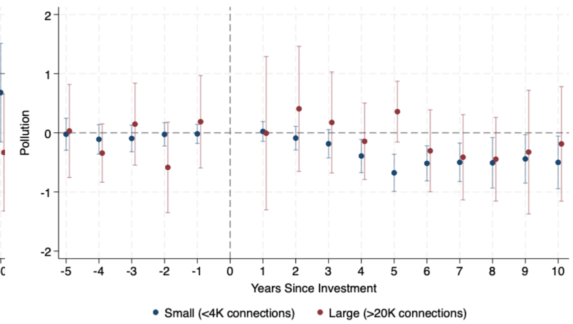
Appendix E. Estimation of Local Political Frictions Parameters

E.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 (l^r, l^q) state space representing rate and expected pollution investment stock variables. These

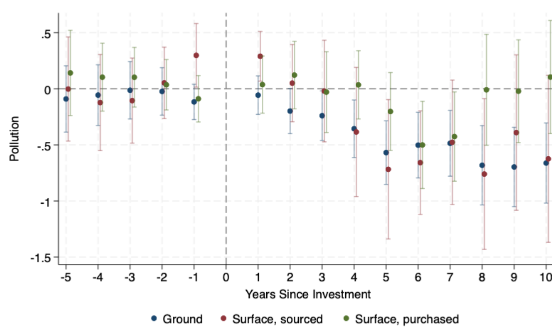
FIGURE D7. Investment Effectiveness Across Water System Types



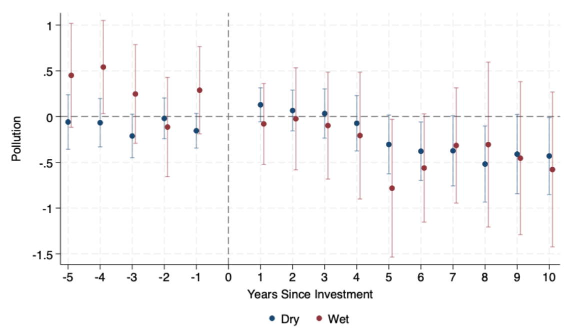
(A) Amount



(B) Size



(C) Water Source



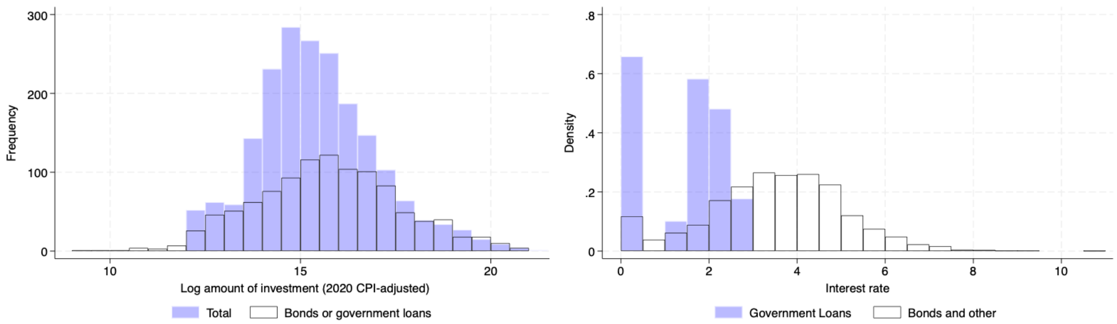
(D) Precipitation

Notes: Each plot presents the coefficient estimates of the dynamic effects of investment on water pollution, using specification (12), 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).

stocks are repleted when investment occurs and depreciate over time to match the shape of our event studies estimated in Figure 3. To reduce the computational complexity, we assume that the governing body does not anticipate future changes in x_{jt}.[1] In each state (l^r, l^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_{jt}, incorporating beliefs over all these states would substantially increase the size of the state space.

FIGURE D8. 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).

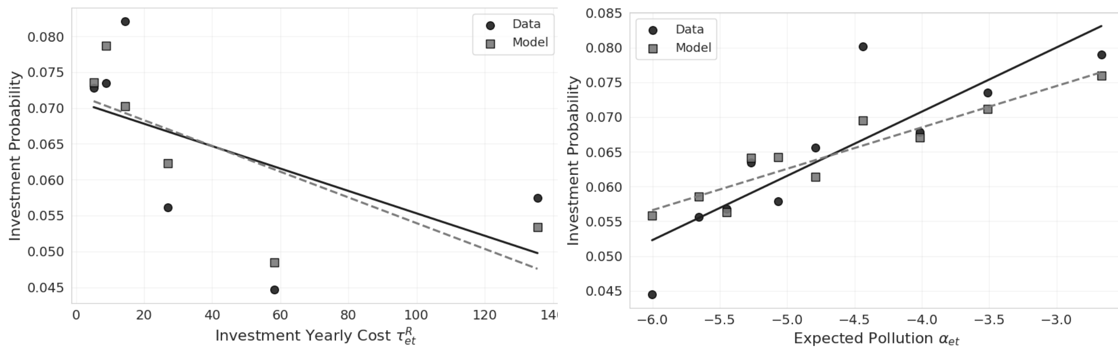
willingness-to-pay $u\big(q_{jt}, r_{jt}; \omega(y_i)\big)$ and aggregate resident utilities into the policy-maker's flow payoff using the income-specific turnout rates:

$$\Pi_{jt}(l^r, l^q; \boldsymbol{\psi}) = \sum_{y} \lambda(x_{jt}; \boldsymbol{\psi})\, \tilde{\pi}_{jt}(y)\, u\big(q_{jt}(l^q), r_{jt}(l^r); \omega(y)\big),$$

where $q_{jt}(l^q)$ and $r_{jt}(l^r)$ denote the mappings from the grid states (l^q, l^r) into current pollution and rates. Note that critically, our measure of accountability A_j is allowed to influence the weight $\lambda(\cdot)$. The law of motion for the discretized state variables implied by our event study estimates are described in Equations 5 and 6. Given these primitives and the private cost of investing $\kappa(x_{jt}; \boldsymbol{\psi})$, 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 $\Delta V_{jt}(l^r, l^q; \boldsymbol{\psi})$ and implied conditional choice probabilities

$$P(a_{jt} = 1 \mid l^r, l^q, x_{jt}; \boldsymbol{\psi}) = \frac{\exp(\Delta V_{jt}(l^r, l^q; \boldsymbol{\psi}))}{1 + \exp(\Delta V_{jt}(l^r, l^q; \boldsymbol{\psi}))},$$

FIGURE E9. 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

$$\ell(\boldsymbol{\psi}) = \sum_{j,t} a_{jt} \log P(a_{jt} = 1 \mid l^r_{jt}, l^q_{jt}, x_{jt}; \boldsymbol{\psi}) \tag{21}$$

$$+ (1 - a_{jt}) \log \big(1 - P(a_{jt} = 1 \mid l^r_{jt}, l^q_{jt}, x_{jt}; \boldsymbol{\psi})\big) \tag{22}$$

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

E.2. **Model fit.** Figure E9 evaluates the model's ability to reproduce key empirical investment patterns. Panel (A) examines the relationship between investment probability and the magnitude of water rate increases associated with investment. In the data, investment becomes less likely as the implied cost of investment rises, reflecting the higher financial burden on residents. The model closely matches this declining relationship, indicating that the estimated cost component of the policymaker's payoff captures the observed sensitivity of investment decisions to financing costs. Panel (B) examines how investment probability varies with expected pollution levels. Both the data and the model show that investment is more likely when expected pollution is higher, reflecting greater expected benefits from improving water quality. Overall, the figure shows that the estimated model successfully reproduces two central empirical patterns: investment is less likely when costs are high and more likely when environmental need is greater.