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Explaining the appearance and success of voter referenda for open-space conservation

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Abstract

This paper provides an empirical investigation of the factors that influence the appearance and success of voter referenda to raise public funds for open-space conservation. We take advantage of a data set that includes detailed information on all such referenda that occurred in the United States between 1998 and 2003. Combining these data with information from the U.S. Census and state-specific variables, we conduct a nationwide analysis along with focused studies of referenda that occurred in New Jersey and Massachusetts. The paper provides the first investigation of how funding mechanisms and funding rates affect voter support for public acquisition of open space. We also provide evidence on the relationship between existing patterns of open space and voter support for open-space referenda. As open-space initiatives continue to gain popularity at the ballot box, the descriptive insights of this paper should prove useful for both policy-makers and advocates working in the area of land use management.

JEL classification: H41; Q38; Q58

Keywords: Open-space; Land conservation; Voter referenda

1. Introduction

The protection of open space from the advance of ‘‘urban sprawl’’ has emerged as one of the more pressing environmental issues in the United States. Open space is generally understood to be a public good that will be under-provided without policy interventions. Policy-makers have begun efforts to protect open space using various instruments—including zoning regulations, development taxes, urban growth boundaries, conservation easements, and public acquisition of undeveloped land. Increasingly, citizens are also becoming directly involved in open-space conservation through ballot initiatives designed to implement mechanisms for public land acquisition. Nearly 1,000 jurisdictions at the state, county, and local levels held open-space referenda

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between 1998 and 2003, and approximately 80% of these initiatives passed, raising over $21 billion for open-space conservation.[1]

The proliferation and high success rate of open-space ballot initiatives raise several economic and policy-relevant questions. What factors contribute to the appearance of an open-space referendum in a jurisdiction? How does an initiative's funding mechanism—such as a bond, property tax, sales tax, or income tax—affect the way citizens vote? How responsive are favorable votes to the costs of an open-space initiative? How do socioeconomic characteristics influence voting results? What is the effect of existing patterns of land use? And what other features of a referendum affect voting outcomes?

These questions motivate our analysis in this paper. We construct a data set of open-space referenda that occurred in the United States between 1998 and 2003. Detailed information on each referendum comes from annual reports, titled *LandVote*, that are published by the Trust for Public Land (TPL) and the Land Trust Alliance (LTA).[2] These data include each referendum's political jurisdiction, proportion voting for and against, financing mechanism, financing rate, land characteristics, and other policy-relevant variables. For each jurisdiction we also collect data from the U.S. Census on socioeconomic characteristics. Then, using the combined data, we estimate econometric models to determine the impact of referendum characteristics and socioeconomic variables on voting results.

In addition to the nationwide analysis, we conduct two focused studies of referenda that occurred in New Jersey and Massachusetts. Statewide policies were passed in both states to provide incentives for local jurisdictions to raise taxes for open-space conservation. The result has been numerous referenda in both states: 237 in New Jersey and 137 in Massachusetts between 1998 and 2003. For both states we collect further Census data on all jurisdictions that did *not* hold a referendum. Additionally, we collect specific data on the amount of open space, recent rates of open-space loss, average property tax burdens, and state-specific features of the referenda. Taking advantage of all these data, we then estimate models for each state in order to determine: (1) what factors influence whether a jurisdiction has held an open-space referendum, and (2) what factors explain the success of a referendum in terms of voting results.

Other researchers have investigated related questions. In a pioneering study of referenda results, Deacon and Shapiro [4] analyze voting outcomes for a law in California to protect coastal zones from development. They find some evidence that the natural coastal environment is a normal good, but the effect is not statistically significant. Kahn and Matsusaka [11] also analyze statewide referenda in California. Three of the referenda they study were to authorize bond issues to purchase park, forest, and wildlife areas. They find evidence that collectively provided open space is a normal good, except when income is very high, in which case it becomes inferior. They also find that people are more likely to vote yes in more urban counties. Another study by Kline and Wichelns [12] uses statewide referenda in Pennsylvania and Rhode Island to investigate demand for the purchase of farmland development rights. They find that the proportion of yes votes increases with a town's population growth, home value appreciation, farmland loss, urbanization, and prevalence of resource sensitive lands.

Because the aforementioned studies use local voting results in statewide referenda, they cannot address the question of what factors contribute to the appearance of an open-space referendum in the first place. Howell–Moroney [9] considers this question in a study of municipalities throughout the Delaware Valley region. He finds that the appearance of a referendum is responsive to patterns of land use, whereby low population density and loss of open space increase the probability of a referendum occurring. He also finds that higher population and median household income increases the probability of a referendum.[3]

[1]Up to date details on the number of open-space referenda and total funds raised can be found on the Trust for Public Land website at http://www.tpl.org.

[2]The reports were published as *Voters Invest in Open Space* between 1998 and 2000 and were renamed *LandVote* beginning in 2001. In total, these reports summarize the results of 968 state, county, and local ballot questions on open space. The reports attempt to be comprehensive and include only referenda involving the direct acquisition of undeveloped land or farmland. Ballot measures for related policies, such as growth controls, are not included.

[3]Howell–Moroney's study is a response to another paper by Romero and Liserio [17]. The latter uses nationwide data on referenda that occurred between 1998 and 1999, and it finds that only socioeconomic factors motivate open-space referenda, while actual patterns of land use do not play a role. Romero and Liserio's conclusion is questionable, however, because of a number of methodological concerns. See Howell–Moroney's paper for a detailed discussion.

Our paper makes four primary contributions to the literature. First, we construct the most comprehensive data set on open-space referenda to date. Second, we take advantage of variation in the financing mechanism across referenda (e.g., bonds or taxes) in order to investigate whether the type of mechanism proposed affects voter support for open-space acquisition. Third, we exploit variation in the funding rates within the different mechanisms (e.g., bond amounts and tax rates) to determine how responsive voters are to the costs of an open-space initiative. Fourth, we conduct detailed analyses of two states in order to determine the factors that influence the appearance of a referendum, in addition to the factors that influence a referendum's success.[4]

The results provide new insights into citizen preferences for open space and the relationship between characteristics of an open-space policy and voter support. We find strong evidence that voters are more likely to approve bonds than tax increases. Not surprisingly, funding rates also matter. Higher rates generally decrease the odds of a yes vote, but interestingly, the opposite result emerges at the state and county levels. In general, we find that the factors influencing referenda outcomes differ between the state–county level and the local level.

In New Jersey and Massachusetts, we find evidence that jurisdictions holding open-space referenda differ significantly from those that do not. Referenda tend to occur in wealthier communities that have experienced greater population growth. While jurisdictions with more open space and recent open-space loss are more likely to have held a referendum, these same variables have different effects on referenda success. In particular, we find that, over certain ranges, more open-space loss in the years prior to a referendum actually reduces voter support. We also find further evidence that collectively provided open space is a normal good. Other findings relate to the importance of farmland as a type of open space, to specific features of the proposed open-space policies, and to the effect of existing tax burdens.[5]

The remainder of the paper is organized as follows. Section 2 describes the data used in the analysis. Section 3 provides details on the econometric specifications. Section 4 reports the results of the nationwide analysis along with the results of the New Jersey and Massachusetts studies. Section 5 concludes with a summary of the main results.

2. Data

We collected data on open-space referenda from the annual *LandVote* survey published by the TPL and the LTA. The *LandVote* survey attempts to provide a comprehensive listing of all open-space referenda that involve the direct acquisition of undeveloped land.[6] Using the information contained in the *LandVote* survey, we generated variables for several characteristics of each open-space referendum. These variables include the referendum's date, whether the initiative passed, proportion voting yes, level of government, funding mechanism, funding rate, whether farmland was included as part of the initiative, and whether the initiative extended an existing program or created a new one.

For each referendum's jurisdiction, we also collected data from the U.S. Census online summary files for 1990 and 2000. These data include each jurisdiction's population, population growth, population density, age profile, household income, home value, and home ownership rate.

Further data were collected for the New Jersey and Massachusetts studies. To compare jurisdictions that did and did not hold a referendum, we collected Census data for *all* local jurisdictions in both states. Additional data were collected on existing property taxes in all local jurisdictions, and on policy variables that are specific to the ballot initiatives that occurred in each state. We also obtained geographic information

[4] Readers should be aware of another study that seeks to make similar contributions. In a recent working paper, Nelson et al. [15] use the same data to analyze the appearance and success of voter referenda in municipal-level, open-space initiatives. They consider a variety of alternative explanatory variables that make for useful comparisons with the results presented here.

[5] Many of our results are comparable to findings in other strands of the literature on open space. These studies employ contingent valuation [2,3,20], stated preferences for different types of open space [13], and revealed preferences for existing open space [1]. We refer to the findings of these studies were appropriate in the discussion of our results.

[6] Publication of the *LandVote* survey began in 1998, and data for selected years are available online at http://www.landvote.org. Data for other years can be obtained from the TPL upon request.

Table 1
Descriptive statistics for open-space referenda

	N	Proportion passing	Proportion yes votes	
			Mean	Std. dev.
Panel A: Summary of referenda results by level of government				
State	23	0.87	0.63	0.10
County	146	0.76	0.58	0.12
Local	688	0.80	0.61	0.12
Total	857	0.80	0.61	0.12

	N	Proportion passing	Proportion yes votes	
			Mean	Std. dev.
Panel B: Summary of referenda results by funding mechanism				
Proptax	299	0.83	0.60	0.10
Proptaxsur	131	0.57	0.51	0.12
Bond	280	0.90	0.66	0.11
Salestax	49	0.65	0.54	0.13
Inctaxsur	28	0.79	0.60	0.11
Other	72	0.79	0.62	0.13

	Units	Mean	Std. Dev.	Min.	Max.
Panel C: Descriptive statistics for funding rates					
Proptaxrate	mills	0.282	0.370	0.001	2.5
Proptaxsurrate	%	2.51	0.78	0.5	3
Bondrate	$10,000,000	4.66	23.50	0.0025	230
Salestaxrate	%	0.437	0.397	0.03	0.2
Inctaxsurrate	%	0.228	0.133	0.00125	0.5

Notes: *Proptax* is property tax, *Proptaxsur* is property tax surcharge, *Bond* is bond, *Salestax* is sales tax, and *Inctaxsur* is income tax surcharge. The funding rate variables correspond with the funding mechanism variables. The total number of referenda in Panel B is 859 because two of the referenda included more than one funding mechanism.

system (GIS) data on existing levels of open space and recent rates of open-space loss. These data and other variables specific to New Jersey and Massachusetts are described in detail in Sections 4.2 and 4.3, respectively.

Of the 968 referenda in the *LandVote* survey between 1998 and 2003, a total of 857 observations were included in the final data set.[7] These referenda occurred in 771 different jurisdictions. Eighty-six jurisdictions held more than one referendum between 1998 and 2003. In some cases more than one attempt was made to pass an open-space policy, while in other cases more than one policy was approved. The entire data set covers 38 different states, although the majority of referenda took place in the northeast and mid-atlantic regions.

Table 1 reports descriptive statistics. Eighty percent of the ballot measures passed, with an average of 61% of the electorate voting yes. A large majority of the referenda were conducted in local jurisdictions (including cities, towns, townships, boroughs, and villages), with comparatively few in counties and states. Property tax increases and bond issues constitute the majority of the funding mechanisms; each accounts for approximately one-third of the measures. Property tax surcharges are the next most prevalent funding mechanism and

[7]The remaining 111 observations were not included for four possible reasons. First, the referenda's jurisdiction was a park district that was not coterminous with any jurisdiction for which we could obtain corresponding Census data (27 observations). Second, even after consulting with TPL staff, the information provided in *LandVote* was not sufficient to match the jurisdiction with a corresponding location in the U.S. Census (14 observations). Third, the initiative passed in a town meeting rather than having been put to a general election (7 observations). Finally, the referendum's financing data was not available because it was missing or the initiative did not involve a direct commitment of funds, as in the case of a simple advisory measure (63 observations).

Table 2
Descriptive statistics for socioeconomics variables

Variable	Definition	Sample	Nation	t-stat.
Population	Population of jurisdiction in 100,000s	–	–	–
Popnchg	Population change between 1990 and 2000 as a proportion	0.28 (1.25)	0.13	3.55
Density	Population density in 1,000s per square mile	1.36 (1.76)	0.10	20.95
Under18	Proportion of population under age 18	0.25 (0.04)	0.26	5.60
Over65	Proportion of population over age 65	0.13 (0.06)	0.12	1.24
Income	Median household income in $10,000s	6.26 (2.10)	4.20	28.83
Homevalue	Median value of owner-occupied housing in $100,000s	2.10 (1.14)	1.20	23.36
Homeown	Proportion of occupied housing units that are owner-occupied	0.76 (0.14)	0.66	20.13

Notes: Data for all variables are from the 2000 U.S. Census, except for *Popnchg*, which is based on data from the 1990 and 2000 Census. Statistics in the Sample column are the unweighted means (standard deviations) for all 857 observations. Statistics are not reported for *Population* because the jurisdictions are not homogenous or comparable for this variable. Statistics in the Nation column are the U.S. Census estimates for the nation as a whole. The national average for *Density* includes only the contiguous 48 states.

constitute 15% of the measures.[8] The remaining referenda are divided among sales tax increases, income tax surcharges, and a category for other funding mechanisms.[9]

It is not reported in Table 1 but worth mentioning that the different financing mechanisms were not evenly distributed among the levels of government. Seventy percent of the state-level referenda were for bond issues, and the remaining 30% were in the "other" category. In contrast, 40% of the county-level referenda were bond issues, and 50% were property tax or sales tax increases. The financing mechanisms were more evenly distributed among the local-level referenda.

Panel B in Table 1 reveals variation in voting outcomes among the financing mechanisms. Bonds generate the highest pass rate and the largest proportion of yes votes. Property tax surcharges and sales taxes, in contrast, generate substantially lower pass rates and proportions of yes votes. These differences suggest that finance mechanisms may affect the outcomes of open-space referenda. In the next section, we specify regression models to test for such mechanism effects while controlling for other factors that may influence election results.[10]

Panel C in Table 1 reports descriptive statistics for the funding rate variables. For example, the mean property tax increase was 0.28 mills (i.e., 28 cents per thousand dollars of tax-assessed value). The variation in magnitude within the funding rates is quite pronounced. The property tax increases, for example, range from 0.001 to 2.5 mills. This difference implies that for a household with a tax-assessed property value of $150,000, the increased tax burden ranges from 15 cents per year to $375 per year. The bond amounts also cover a wide range, from $25,000 in Baltimore County, Maryland to $2.3 billion in the state of California. With our econometric analysis, which we describe in the next section, we also test whether funding rates affect voting outcomes.

The socioeconomic variables that we obtained from the U.S. Census are defined in Table 2. When appropriate, we report means for the sample of jurisdictions that held a referendum and compare them to the national averages. The t-statistics in Table 2 are based on a test of whether the sample mean is statistically different from the national average. All tests are statistically significant except for the one comparing the proportions of the population over the age of 65. Compared to national averages, jurisdictions that have held

[8]A property tax surcharge imposes a percentage increase in one's property tax bill. A surcharge is distinct from a property tax millage because it typically allows for additional exemptions and/or may be levied on a subset of the taxable population. The units also differ: property taxes are based on a millage applied to tax-assessed value, and surcharges are based on a percentage applied to a property-tax bill.

[9]The *Other* category includes parcel taxes, real estate transfer taxes, use taxes, retailers' occupation taxes, lottery taxes, hotel taxes, and intragovernmental transfers.

[10]A distinct advantage of the data is the fact that jurisdictions are often constrained regarding their choice of funding mechanism. In New Jersey and Massachusetts, for example, state law requires property taxes and property tax surcharges, respectively. Thus, in our econometric models, it is reasonable to treat the funding mechanism as exogenous to the level of voter support.

an open-space referendum tend to have higher population growth and greater density.[11] They also tend to have greater household incomes, home values, and home ownership rates. In our studies of New Jersey and Massachusetts, we estimate probit models to test more formally for differences between those jurisdictions that have held a referendum and those that have not.

3. Econometric specifications

We estimate regression models to explain the election outcomes of open-space referenda. The dependent variable in our models is

$$logodds_i = \ln\left(\frac{P_i}{1-P_i}\right), \tag{1}$$

where P_i is the proportion of yes votes out of the total number of votes cast in referendum i. This variable is the log-odds ratio, and it is commonly used in econometric models of aggregate voting results [e.g., 4,5,11,12,18,20].

The equations that we estimate for the nationwide data have the general form:

$$\begin{aligned} logodds_i = {} & \boldsymbol{\beta}_1' \boldsymbol{Mech}_i + \boldsymbol{\beta}_2' \boldsymbol{Rate}_i + \beta_3 Extend_i + \beta_4 Farm_i + \beta_5 NotNov_i \\ & + \boldsymbol{\beta}_6' \boldsymbol{Socio}_i + \boldsymbol{\beta}_7' \boldsymbol{Gov}_i + \delta_s + \lambda_t + \varepsilon_i, \end{aligned} \tag{2}$$

where $\boldsymbol{Mech}_i$ is a categorical variable indicating the referendum's funding mechanism; $\boldsymbol{Rate}_i$ is a vector of the funding rate variables that equal zero if the funding rate does not apply; $Extend_i$ is a dummy variable indicating whether the referendum extends an existing policy; $Farm_i$ is a dummy variable indicating whether farmland is part of the proposed land acquisitions; $NotNov_i$ is a dummy variable indicating whether the referendum took place off a typical election cycle in November; $\boldsymbol{Socio}_i$ is a vector of the socioeconomic variables that includes those listed in Table 2; $\boldsymbol{Gov}_i$ is a categorical variable indicating whether the referendum occurred at the state, county, or local level; δ_s is a state-specific intercept; λ_t is a year-specific intercept; and ε_i is a random error term.

The log-odds model specified in (2) has a microeconomic foundation. Deacon and Shapiro [4] develop a model that begins with individual preferences and aggregates up to collective voting results. The log-odds specification is a simplified version of their model's empirical implication.[12] This micro foundation implies that the aggregate voting results can be used to make inferences about voter preferences. The validity of making such inferences has empirical support as well. A study by Fischel [6] found little difference in a comparison between aggregate voting results and individual preferences for an environmental referendum in New Hampshire.

The independent variables in regression model (2) are useful for answering several questions and controlling for potentially important effects. The inclusion of $\boldsymbol{Mech}_i$ enables consideration of how voting outcomes respond to the funding mechanism of the initiative. To accurately test for such differences, all of the continuous variables in Eq. (2), which include $\boldsymbol{Rate}_i$ and $\boldsymbol{Socio}_i$, are demeaned.[13] The coefficients on $\boldsymbol{Rate}_i$ will provide an estimate of the responsiveness of voter support to the funding rates. $Extend_i$ and $Farm_i$ will determine whether selected characteristics of the open-space proposal affect its success. $NotNov_i$ will determine whether holding the referendum outside the typical cycle in November affects voter support. $\boldsymbol{Socio}_i$ will be useful to detect factors that influence preferences for open space, such as income and population density. $\boldsymbol{Gov}_i$

[11] It is worth keeping in mind that the national average for *Density* is based on all land area in the lower 48 states, including the large and sparsely populated areas in the west. The difference in Table 2 with respect to *Density* may therefore be misleading.

[12] The simplification arises because, like most studies, we do not consider abstentions. Rubinfeld [18] develops a voting model that ignores abstentions and also generates a log-odds specification. While his model is discussed in the context of voting in local school elections, the analysis can apply equally to open-space measures.

[13] The demeaning implies that estimated coefficients on $\boldsymbol{Mech}_i$ are interpreted as differences between the funding mechanisms based on the average funding rate within each mechanism. The demeaning avoids potentially erroneous conclusions that mechanism X is preferred to Y, when in fact the differences are due to X having more favorable rates than Y. The variables in $\boldsymbol{Socio}_i$, which are also continuous, are demeaned for similar reasons—to estimate coefficients on $\boldsymbol{Mech}_i$ that are not biased by potential correlation between the funding mechanisms and the socioeconomic variables.

will indicate whether election results differ between levels of government. Finally, the state- and year-specific intercepts will control for unobserved state and year heterogeneity.

We estimate the models using weighted least squares (WLS) to account for heteroskedasticity due to the analysis of averaged, grouped data. The weight for each observation i is $(n_i\hat{P}_i(1-\hat{P}_i))^{1/2}$, where n_i is *Population* and $\hat{P}_i$ is the predicted proportion of yes votes.[14] Using these weights implements the minimum chi-squared estimator, and the effect (which we discuss more in the next section) is to place more weight on referenda in jurisdictions with larger populations.

Our analysis of New Jersey and Massachusetts referenda differs somewhat because of the availability and nature of the data. Using the additional Census data on jurisdictions that did not hold an open-space referendum, we are able to investigate the factors that influence the appearance of a referendum in a local jurisdiction. We assume the underlying propensity of a jurisdiction to hold a referendum is given by

$$r_i^* = \gamma_0 + \gamma_1' \boldsymbol{Socio}_i + \gamma_2 Avproptax_i + \gamma_3' \boldsymbol{OSvars}_i + u_i, \tag{3}$$

where $Avproptax_i$ is a variable for the jurisdiction's average residential property tax payment, and $\boldsymbol{OSvars}_i$ is a vector of variables characterizing open space in the jurisdiction (including existing levels of open space and recent rates of open-space loss).[15] While r_i^* is unobservable, we do observe whether or not a jurisdiction actually held a referendum. The observations can therefore be written as $r_i = 1$ if $r_i^* > 0$, or $r_i = 0$ if $r_i^* \leqslant 0$. Assuming u_i is normally distributed, the parameters of (3) can be estimated with a probit model:

$$\Pr(r_i = 1) = \Phi(\gamma_0 + \gamma_1' \boldsymbol{Socio}_i + \gamma_2 Avproptax_i + \gamma_3' \boldsymbol{OSvars}_i), \tag{4}$$

where $\Phi(\cdot)$ is the standard normal cumulative distribution function. We estimate this model for both New Jersey and Massachusetts in order to determine how the explanatory variables affect a jurisdiction's probability of holding an open-space referendum.

We also estimate log-odds models for New Jersey and Massachusetts to explain election outcomes where they did occur. For these models, we include only those referenda that were part of statewide initiatives to encourage open-space conservation (described in Sections 4.2 and 4.3). These include all of the 237 referenda in New Jersey and 122 of the 137 referenda in Massachusetts. A nice feature of these data is that all referenda within each state proposed the same funding mechanism but with varying funding rates. Accordingly, the models require fewer explanatory variables than were included in (2). The estimated equations have the general form:

$$logodds_i = \psi_1 Rate_i + \boldsymbol{\psi}_2' \boldsymbol{X}_i + \boldsymbol{\psi}_3' \boldsymbol{Socio}_i + \psi_4 Avproptax_i + \boldsymbol{\psi}_5' \boldsymbol{OSvars}_i + \theta_t + v_i, \tag{5}$$

where $Rate_i$ is a variable for the funding rate of the mechanism within the state, $\boldsymbol{X}_i$ is vector of state-specific policy variables, θ_t is a year-specific intercept, and v_i is a random error term. These models are estimated with WLS using the same weights as those discussed previously.[16]

For each state we also estimated Heckman two-step models linking Eqs. (4) and (5). This procedure accounts for potential sample-selection bias that may arise from estimating Eq. (5) on its own. The approach, however, is somewhat limited in this context because of the nature of the data. To fully correct for sample-selection bias, all of the variables in the regression (5) would need to be in the selection Eq. (4) as well (see Greene [7]).[17] But this is not possible here, as data for $Rate_i$ and $\boldsymbol{X}_i$ only exist for places that actually held a referendum. We nevertheless estimate the selection models to account for some degree of potential bias, but as we discuss later, the selection bias is not statistically significant in any of the estimated equations.

[14] We use ordinary least squares to obtain initial values of $\hat{P}_i$. We then iterate to convergence with WLS. See Greene [7] for details on the use of these weights and on the method of estimation.

[15] We describe in detail the data used for these specific variables in Sections 4.2 and 4.3.

[16] We collected data for other variables that are not included in the final specifications because the results were never statistically significant. For both states, these include variables based on voting results for the 2000 presidential election. For Massachusetts, we also found statistically insignificant effects for the mechanism by which the referendum was put on the ballot (see footnote 41 for more details).

[17] Ideally, in order to avoid identification on functional form alone, there should also be an identifying restriction of at least one variable in Eq. (4) that is not in Eq. (5).

4. Estimation and results

We report the econometric results in this section. Those for the nationwide analysis are reported first, followed by those for New Jersey and Massachusetts.

4.1. Nationwide

We begin with estimates of Eq. (2) using the nationwide sample of 857 referenda. These results are reported as the pooled model in Table 3. We also estimate Eq. (2) using two subsets of the data: one includes all of the local-level referenda, and the other includes all of the state- and county-level referenda. These models are reported as the local and state–county models in Table 3.[18] For all three models, we report standard errors that are robust to clustering at the jurisdiction level; this accounts for the fact some referenda occurred within the same jurisdiction and therefore may not be entirely independent observations.[19]

The reason for estimating the pooled and separate equations is to test whether the explanatory variables affect local results differently than state and county results. Combining all the data may be overly restrictive because of the differences in scale and political dynamics between these levels of government. One might expect, for example, that demand for open space may differ at the local level because of the smaller number of taxpayers and/or the closer proximity to the proposed land acquisitions.[20]A further reason for splitting the data in this manner follows from the weighted estimation, which places more weight on the higher-population (i.e., state and county) observations in the pooled model. This explains why many of the coefficient estimates in the pooled model are closer to those in the state–county model than they are to those in the local model. To test formally for differences, we conduct a heteroskedasticity-robust F-test comparing the separate estimates to the pooled estimates. The results indicate statistically significant differences between the local and state–county results ($F[21, 770] = 4.87, p < 0.01$). In the following discussion, therefore, we focus primarily on the results of the local and state–county models.

First consider the effects of the different financing mechanisms. Bonds are the omitted category, so coefficients are interpreted as the pairwise comparison between the indicated mechanism and bonds, evaluated at the mean values for each of the funding mechanisms and the socioeconomic variables (due to the demeaning described previously). A clear pattern emerges from the results: nearly all coefficients are negative, indicating that voters are more likely to approve bond issues than tax increases. The difference between sales taxes and bonds is statistically significant in all models. In the local model, all of the specific mechanisms are significantly different from bonds. In the state–county model, *Proptax* is the only mechanism that is not significantly different from bonds.

Several coefficients on the statistically significant funding mechanisms have magnitudes close to (or in excess of) -0.40, which implies a decrease of approximately 33% on the odds ratio.[21] This implies that, beginning from an average of 60% of the electorate voting yes (with an odds ratio of 1.5), the model predicts that switching from a bond to one of the taxes would decrease the proportion of yes votes to roughly 50% (with an odds ratio of 1). Note that this difference can be pivotal for a referendum that requires a majority to pass. In such cases, therefore, financing with a bond versus a tax can make the difference between whether or not an open-space referendum passes.

So why might voters prefer bonds? We suggest four possible reasons. First, citizens may perceive bonds to generate the necessary funding for an acquisition immediately, while they may not expect benefits from a tax for several years. Second, bonds are more likely to be associated with specific open-space acquisitions, whereas tax revenues are more likely to accrue in a fund with unspecific future benefits. Third, the costs of a bond may

[18]In addition to the results reported in Table 3, we estimated each model without state fixed effects, but the results were very similar.

[19]Throughout the paper, clustered standard errors refer to the Huber/White/sandwich estimator of the covariance matrix that is generalized to account for arbitrary correlation among the error terms within the same jurisdiction (cluster). See Williams [21] for a description of the estimator and Rogers [16] for details on its implementation using Stata. While the clustering has no effect on the coefficient estimates, it produces standard errors that are generally larger than those produced without clustering.

[20]Breffle et al. [2] and Champ et al. [3] find empirical evidence that people place greater value on open space when they live closer to it.

[21]The proportional change in the odds ratio with a change in a dummy variable from 0 to 1 is given by $e^{\beta} - 1$, where β is the coefficient on the dummy variable [8].

Table 3
Nationwide, local, and state–county WLS regressions

	(1) Pooled		(2) Local		(3) State–county	
	Coef.	*t*-stat.	Coef.	*t*-stat.	Coef.	*t*-stat.
Proptax	-0.302^{***}	(3.35)	-0.293^{*}	(1.83)	-0.130	(1.01)
Proptaxsur	-1.026^{***}	(3.12)	-0.724^{**}	(2.36)	–	–
Salestax	-0.508^{***}	(4.15)	-0.572^{***}	(3.79)	-0.528^{***}	(3.28)
Inctaxsur	-0.902^{***}	(4.42)	-0.394^{**}	(2.45)	–	–
Other	-0.320^{***}	(3.28)	0.040	(0.34)	-0.291^{**}	(2.42)
Proptaxrate	0.366^{**}	(2.41)	-0.173	(1.05)	0.497^{**}	(2.32)
Proptaxsurrate	-0.088	(1.03)	-0.054	(0.67)	–	–
Bondrate	0.001^{**}	(1.99)	-0.041^{***}	(3.09)	0.002^{**}	(2.09)
Salestaxrate	-0.190	(0.88)	0.069	(0.49)	-0.091	(0.30)
Inctaxsurrate	-1.078	(1.15)	-1.035	(1.23)	–	–
Extend	0.561^{***}	(4.68)	0.140	(0.92)	0.511^{***}	(3.55)
Farm	-0.147^{**}	(2.07)	0.162^{*}	(1.90)	-0.152	(1.62)
NotNov	0.190^{***}	(2.81)	0.257^{**}	(2.28)	0.123	(1.48)
Population	-0.001	(1.37)	0.042^{***}	(4.29)	-0.000	(0.27)
Popnchg	-0.019	(0.85)	0.010	(0.68)	-0.163	(0.67)
Density	0.011	(0.55)	-0.008	(0.36)	-0.002	(0.05)
Under18	1.237	(0.93)	-2.418^{**}	(2.37)	1.461	(0.48)
Over65	1.494^{*}	(1.89)	0.164	(0.27)	1.944	(0.94)
Income	0.073^{***}	(2.82)	0.061^{***}	(3.18)	0.121	(1.58)
Homeown	-1.196^{**}	(2.37)	-0.376	(0.82)	-1.111	(0.84)
State	-0.025	(0.21)	–	–	0.018	(0.10)
County	-0.097	(1.40)	–	–	–	–
State dummies	Yes		Yes		Yes	
Year dummies	Yes		Yes		Yes	
Observations	857		688		169	
R-squared	0.79		0.56		0.89	

Notes: The dependent variable in all models is *logodds*. *Bond* is the omitted category for the funding mechanisms. *Local* is the omitted category for government level in the pooled model. Significance levels are based on standard errors that are robust to clustering at the Jurisdiction level. One, two, or three asterisks indicate significance at the levels $p<0.10$, $p<0.05$, or $p<0.01$, respectively.

be delayed compared to the immediate costs of a tax increase. Fourth, citizens may not have a clear idea about the costs of bonds, whereas the costs of tax increases are readily apparent. Note that each of these possible explanations conflicts with the general notion of Ricardian equivalence, which implies that citizens should be indifferent between the different funding mechanisms. Thus, we conclude that either Ricardian equivalence does not hold in this context, or that voters are subject to a form of fiscal illusion.[22]

Before looking at the results for the funding rates, it is important to recognize that standard price effects do not apply. While one might expect a higher funding rate to decrease support for an initiative, this need not be the case. The reason is that open-space acquisitions are endogenous to the funding level—that is, higher rates enable the purchase of more (or higher valued) land. Voters must therefore consider two effects that occur with an increase in the funding rate. One effect is having to pay more oneself. The other is enjoying the additional (or higher valued) open space generated by revenues from all taxpayers. The former effect can decrease a voter's welfare, while the latter effect increases it. Thus, the sign of the coefficients on the funding rates can be interpreted as an indicator of which effect dominates. Negative coefficients would suggest that the

[22]The different ideas suggested in this paragraph are working hypotheses. We are unable to distinguish between them in this study, yet they all underscore the importance of characterizing empirical trends in local public finance and beginning to think about theoretical underpinnings. As such, work of this type contributes to the emerging agenda in the new field of behavioral public finance [14].

effect of having to pay more oneself dominates; positive coefficients would suggest that the "spillin" effect of greater revenues dominates.[23]

The results provide evidence in both directions. In the local model, all funding rate coefficients, except for the one on sales taxes, are negative, suggesting that higher funding rates decrease voter support at the local level. But the only coefficient that is statistically significant is the one on bond rates. The results are different in the state–county model. The coefficients on property tax and bond rates are both positive and statistically significant. These differing results in the local and state–county models can be explained with the two countervailing effects of having to pay more oneself versus benefiting from spillins.

First consider a tax. A citizen's personal tax burden does not depend on the political jurisdiction—for example, a property tax increase of 1 mill imposes the same cost regardless of whether a state, county, or town collects the revenue. In contrast, one would expect the spillins to be smaller in local jurisdictions than in states and counties. Thus, it would not be surprising for the effect of the individual tax burden to dominate the spillin effect at the local level, but not at the state or county level. This reasoning is consistent with *Proptaxrate* having a negative (though insignificant) effect in the local model and a positive effect in the state–county model.

Now consider how the effect of the bond rate may differ between levels of government. In this case, a voter's personal cost of a bond does in fact depend on the political jurisdiction because more people share the costs of a bond in a state or county than in a local jurisdiction. In contrast, spillins for a given bond rate will not depend on the size of the jurisdiction, assuming that the open-space acquisition is a public good. With bonds, therefore, one might expect the individual burden effect to dominate at the local level, but not at the state and county level. The results follow this pattern, as the coefficient on *Bondrate* is negative in the local model but positive in the state–county model.

Another factor that may influence voting outcomes is whether the referendum extends an existing policy or initiates a new one. The results provide some evidence that voters are more likely to reauthorize an existing open-space policy. The coefficient on *Extend* is positive in all three models and statistically significant in the pooled and state–county models. The magnitude of the coefficient in the state–county model implies that, starting from 60% of the voters voting yes, having the initiative be an extension increases the percent voting yes to 72%—a substantial increase. This result is intuitive because jurisdictions with extensions have already revealed a preference and willingness to pay for open space.

Two other variables related to referendum characteristics are statistically significant in the local model. The coefficient on *Farm* is positive, indicting that voters are more likely to favor local farmland preservation than non-agricultural open space. This result corroborates Kline and Wichelns' [13] finding that individuals prefer farmland to most other types of open space. The coefficient on *NotNov* is also positive, implying that voters are more likely to support referenda that occur outside the typical election cycle in November. This result suggests that voter turnout is more favorable for open-space initiatives that are not on the ballot in November. The result may also be capturing more subtle effects related to the interaction among multiple items on a ballot that may or may not be included in November.

Fewer of the socioeconomic variables are statistically significant. In the local model, the coefficient on *Population* is positive and significant, indicating that voters in local jurisdictions with larger populations are more likely to vote yes, due possibly to the spillin effects discussed above. The negative coefficient on *Under18* implies that a greater proportion of the population under 18 years of age decreases voter support. The coefficient on *Income* is also positive and significant, suggesting that publicly provided open space is a normal good. Other studies have found mixed results for the income effect: some find evidence that open space is a normal good [1,2], some find no significant effect [4,12,17], and one study finds that open space is generally a normal good but may become inferior at high levels of income [11].[24] None of the socioeconomic variables has statistically significant explanatory power in the state–county model.[25]

[23]In the public finance literature, "spillins" refer to the public-good benefits that individuals receive beyond what they actually pay for—that is, the benefits they enjoy from the public-good provision by others.

[24]We attempted to replicate Kahn and Matsusaka's [11] result by estimating all of the models with the inclusion of a quadratic term for median household income. While the same pattern emerged, neither of the income coefficients were statistically significant.

[25]We do not include *Homevalue* in any of the regressions in Table 3 because it is highly correlated with *Income*. Inclusion of both variables renders neither statistically significant in all models. This is not the case, however, in the New Jersey and Massachusetts models.

While the nationwide analysis illuminates some of the factors that influence voter support for open-space referenda, it does not shed light on why these referenda occur in the first place. An open-space referendum is not a random event, but an outgrowth of economic, political, and environmental factors that motivate citizens and lawmakers to put open-space initiatives on the ballot. In the following studies of New Jersey and Massachusetts, we address the question of what factors influence the occurrence of an open-space referendum, in addition to the question of what factors influence voting results.

4.2. New Jersey

New Jersey is the most highly represented state in the data set. Since 1989 state legislation has been in place that enables local jurisdictions to impose property taxes for the purpose of open-space acquisition. The state approved further legislation in 1997, called the Green Acres Planning Incentive Program, to provide matching funds to municipalities that adopt open-space property taxes. In order to encourage immediate acquisition of open space, an additional provision of the legislation was that communities can receive 2% interest loans from the state upon approval of a property tax increase.

Between 1998 and 2003, a total of 237 property-tax referenda took place in 178 different local jurisdictions in New Jersey. Fifty-nine of these referenda occurred in jurisdictions that held at least one prior referendum over the same period. We investigate all of the New Jersey election outcomes, along with the additional question of what factors contribute to the appearance of a referendum in a jurisdiction. To accomplish this, we collected data for all 566 local jurisdictions in New Jersey in order to estimate the probit model specified in Eq. (4). The dependent variable equals 1 if the jurisdiction held at least one open-space referendum between 1998 and 2003, and it equals 0 otherwise.

Three new variables are included in all of the New Jersey models. First, *Avproptax* is each jurisdiction's average, annual residential property tax payment in 2000.[26] Second, *Openspace* is the percentage of undeveloped land in 1995 within each jurisdiction's county.[27] Third, *OSloss* is the percentage change in undeveloped land within each jurisdiction's county between the years 1986 and 1995.

The estimated probit model is reported in the first column of Table 4. The results reveal several differences between jurisdictions that held open-space referenda and those that did not. Jurisdictions with higher populations were more likely to hold a referendum. Higher population growth between 1990 and 2000 also raises the likelihood of an open-space initiative. This result is not surprising to the extent that greater population growth is likely to make growth management issues more salient within a community.[28]

More open space in a jurisdiction's county—measured by undeveloped land as a percentage of total land area—increases the probability of a referendum appearing on a ballot. One possible interpretation of this result is Tiebout sorting, whereby people who value open space highly choose to live in places with more open space and are also more likely to support an open-space ballot initiative. Not surprisingly, more open-space loss between 1986 and 1995 also increases the likelihood of a referendum. The magnitude of the coefficient on *OSloss* implies that, starting from the mean *OSloss* of 5.1%, a 1% increase in open-space loss increases the probability of a referendum occurring by 2.8%.[29]

Other results from the probit model are that income and home value are both statistically significant, but have opposite signs. Higher income jurisdictions are more likely to hold an open-space referenda, and jurisdictions with higher home values are less likely to hold an open-space referendum. These results are consistent with open space being a normal good and higher home values implying a greater tax burden for a given property tax increase. We also find that greater home ownership increases the probability of a referendum, perhaps reflecting a greater incentive to capitalize the benefits of open space in home values.[30]

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[26]We obtained this data from the New Jersey Division of Local Government Services, and it is available online at http://www.state.nj.us/dca/lgs/taxes/taxmenu.shtml.

[27]This variable is based on data provided by the New Jersey Department of Environmental Protection. The data is available only at the county level and for the years 1986 and 1995. The data files are available online at http://www.state.nj.us/dep/gis/.

[28]We do not include *Density* as an explanatory variable when data on open space is available.

[29]This result for open-space loss, along with those for population and population growth, are consistent with Howell–Moroney's [9] findings in the Delaware Valley region.

[30]Hedonic studies have found evidence that open space has a positive effect on residential property values [10,19].

Table 4
New Jersey probit and WLS regressions

	(1) Probit		(2) WLS All Obs.		(3) WLS No Repeat	
	Coef.	*t*-stat.	Coef.	*t*-stat.	Coef.	*t*-stat.
Proptaxrate	–	–	−0.566	(1.56)	−0.794**	(2.56)
Extend	–	–	0.039	(0.41)	−0.025	(0.09)
Farm	–	–	0.192**	(2.21)	0.325***	(4.03)
Repeat	–	–	−0.402***	(3.66)	–	–
Priorpass	–	–	0.404***	(2.76)	–	–
Population	1.119***	(4.01)	−0.136	(0.76)	−0.068	(0.49)
Popnchg	0.299**	(2.24)	0.006	(0.12)	0.168	(1.62)
Openspace	0.008**	(2.28)	−0.003	(1.05)	−0.002	(0.83)
OSloss	0.086**	(2.53)	−0.048**	(1.99)	−0.047**	(2.48)
Under18	−0.574	(0.28)	−0.561	(0.31)	−0.524	(0.34)
Over65	−0.116	(0.07)	1.539*	(1.80)	1.570*	(1.79)
Income	0.194**	(2.23)	0.117	(1.55)	0.124*	(1.87)
Homevalue	−0.239**	(1.98)	−0.207	(1.45)	−0.315**	(2.49)
Homeown	2.280***	(3.27)	−0.352	(0.62)	0.014	(0.03)
Avproptax	0.056	(0.91)	−0.010	(0.20)	0.067	(1.24)
Year dummies	–		Yes		Yes	
Observations	566		227		170	
Log likelihood	−293.96		–		–	
% predicted	70.49		–		–	
R-squared	–		0.35		0.39	

Notes: The dependent variable in the probit model is equal to 1 if the jurisdiction ever held a referendum and 0 otherwise. The dependent variable in the WLS model is *logodds*. Significance levels in column (2) are based on standard errors that are robust to clustering at the Jurisdiction level. One, two, or three asterisks indicate significance at the levels $p<0.10$, $p<0.05$, or $p<0.01$, respectively.

We now turn to the WLS estimates of Eq. (5). Mirroring the nationwide analysis, we include the variables *Proptaxrate*, *Extend*, and *Farm*. Because of the relatively high proportion of jurisdictions that held more than one referendum, we include two additional variables to investigate the interaction between repeat initiatives. *Repeat* is a dummy variable indicating whether the jurisdiction held one or more prior open-space referenda within the study period. *Priorpass* is a dummy variable indicating whether one or more of the prior referenda passed. We estimate the model using all of the New Jersey referenda that were part of the Green Acres Program.[31] These results are reported in the second column of Table 4. Once again, the reported standard errors are robust to clustering at the jurisdiction level. For purposes of comparison, we also estimate the model excluding referenda that were repeat initiatives within a jurisdiction. While this model is based on fewer observations, it enables us to focus on first-time referenda and to report unclustered standard errors that are not biased. These results are reported in the third column of Table 4.[32]

Both models generate similar results, although differ somewhat with respect to the coefficients that are statistically significant. The effect of *Proptaxrate* is negative, but only significant in the model without repeat observations. For first-time referenda, the magnitude of the coefficient implies that an increase in the property tax rate of 0.1 mills decreases the proportion of yes votes from 60% to 58% on average.

The negative and statistically significant coefficient on *Repeat* indicates that voters are less supportive of subsequent ballot initiatives in their jurisdiction. Nevertheless, the positive and statistically significant

[31]This includes all 237 referenda that occurred in New Jersey; however, only 227 observations are used in the estimation. The reason for the difference is that data for *Proptaxrate* was missing for 10 observations.

[32]As discussed in Section 3, we also estimated two-step sample selection models using the probit model in the first stage and a WLS regression in the second. Even after attempting various identification restrictions in the first stage, we never found significant sample-selection bias. In all cases, the coefficient on the inverse Mills ratio was statistically insignificant, and the point estimates on other variables did not change in meaningful ways. We therefore report only the WLS regressions on the basis that the sample selection was never significant, and that the inclusion of the inverse mills ratio results in less precise estimates due to collinearity with variables included in both stages (see Wooldridge [22]).

coefficient on *Priorpass* indicates that if a prior initiative passed, the effect of *Repeat* is attenuated. In fact, the two effects are statistically offsetting according to a test of whether the two coefficients sum to zero ($t = 0.01,\ p = 0.98$). Thus, voting outcomes for referenda in jurisdictions that have already passed an open-space initiative are statistically indistinguishable from those having never held a referendum.[33]

New Jersey voters are more likely to vote in favor of local farmland preservation than non-agricultural open space. In both models, the coefficient on *Farm* is positive and statistically significant. This result is similar to that found in the nationwide analysis.

The coefficients on the open-space variables are somewhat surprising. In contrast to the probit model, the coefficient on *Openspace* is not statistically significant in either of the WLS regressions. But even more surprisingly, the coefficient on *OSloss* is negative and statistically significant in both models. This implies that more open-space loss in a local jurisdiction's county decreases voter support for an open-space initiative, although the magnitude of the effect is very small: a 1% increase in open-space loss reduces voter support from an average of 60% to 59%.

To investigate this relationship further, we reestimated both models with a quadratic term for *OSloss*. The coefficient estimates on the *OSloss* variables from both sets of data are identical to the second decimal place:

$$\underset{(1.37,\ 1.69)}{0.21} \times OSloss - \underset{(1.68,\ 2.09)}{0.02} \times OSloss^2,$$

where the terms in parentheses are t-statistics for the estimates corresponding to the models in columns (2) and (3), respectively.[34] These results suggest, although rather weakly, that voter support increases at relatively low levels of open-space loss, then begins to decrease with greater levels of open-space loss. The turning point at 5.25% occurs well within the range of the data for jurisdictions that held a referendum, which has a median of 5.1%, a min of 1.6, and a max of 8.1. Admittedly, it is difficult to explain what may be driving these results, but the fact that we have data for *OSloss* at the county level rather than the local level is a limitation. In our study of Massachusetts, we address this limitation because open-space data are available at the local level. We discuss this further in Section 4.3, where for Massachusetts we also find a negative linear relationship with *OSloss*, but the coefficients in the quadratic specification have the opposite signs.

Three of the socioeconomic variables are statistically significant in models (2) and (3) in Table 4. The positive coefficient on *Over65* implies that a higher proportion of senior citizens in a jurisdiction increases voter support. In the No Repeat model, the effects of *Income* and *Homevalue* are statistically significant and follow the same pattern as that in the probit model. Greater income increases voter support in addition to the probability of a referendum occurring. In contrast, greater home value decreases voter support in addition to the probability of a referendum occurring. In both models, the effect of the average, residential property tax payment is statistically insignificant.

4.3. Massachusetts

The Massachusetts Community Preservation Act (CPA) was passed in 2000. The law is similar to New Jersey's Green Acres program in that it offers state matching funds to communities that raise property taxes for open-space conservation.[35] Rather than levying a property tax millage, however, the CPA authorizes communities to levy a surcharge of up to 3% on existing property tax bills. Optional provisions of the policy include three exemptions from the surcharge: one for low-income families and low- to moderate-income senior citizens; one for the first $100,000 of the tax-assessed value of all properties; and one for commercial and industrial properties.

[33]The variable *NotNov* is not included in the New Jersey models because only one referendum occurred outside of the November election cycle.

[34]We report only the results for the *OSloss* variables because the coefficients on all of the other variables were nearly identical.

[35]In addition to promoting open space, the CPA is intended to promote historic preservation and affordable housing. The law requires that at least 10% of the funds be spent on each of the three objectives, with the remaining 70% allocated at the local legislature's discretion. Specific allocations were not specified prior to elections. To date, approximately 42% of CPA-funded projects have been for open space and recreation. Up to date details are available on the Massachusetts Community Preservation Coalition's webpage at http://www.communitypreservation.org.

Table 5
Massachusetts probit and WLS regressions

	(1) Probit		(2) WLS	
	Coef.	*t*-stat.	Coef.	*t*-stat.
Proptaxsurrate	–	–	−0.238***	(3.83)
Lowinc	–	–	0.263**	(2.40)
First100K	–	–	−0.106	(0.85)
Comind	–	–	−0.138	(1.04)
Repeat	–	–	−0.250	(1.19)
NotNov	–	–	0.128	(1.08)
Population	0.813	(1.48)	−0.053	(1.38)
Popnchg	1.199	(1.64)	0.192	(0.36)
Openspace	0.002	(0.33)	0.007**	(2.02)
OSloss	0.032**	(2.16)	−0.016*	(1.81)
Under18	−11.200***	(3.44)	−3.605	(1.63)
Over65	−1.165	(0.46)	−3.601**	(2.35)
Income	−0.033	(0.27)	0.064	(0.84)
Homevalue	0.211	(1.44)	0.249**	(2.25)
Homeown	2.331**	(2.06)	−0.143	(0.16)
Avproptax	0.373**	(2.29)	−0.174**	(2.09)
Year dummies	–		Yes	
Observations	351		122	
Log likelihood	−192.16		–	
% predicted	70.37		–	
R-squared	–		0.47	

Notes: The dependent variable in the probit model is equal to 1 if the jurisdiction ever held a referendum and 0 otherwise. The dependent variable in the WLS model is *logodds*. Significance levels in the WLS model are based on standard errors that are robust to clustering at the Jurisdiction level. One, two, or three asterisks indicate significance at the levels $p<0.10$, $p<0.05$, or $p<0.01$, respectively.

Our analysis of referenda that were part of the CPA follows the same methodology as that for New Jersey. We collected socioeconomic data for all 351 local jurisdictions in Massachusetts, of which 115 held a CPA referendum.[36] In order to account for the different exemptions, we collected further data from the Massachusetts Community Preservation Coalition on which exemptions applied to each referendum. From this data we generated three dummy variables—*Lowinc*, *First100K*, and *Comind*—to indicate whether the respective exemption applied. The variable *Avproptax* is still defined as each jurisdiction's average, annual residential property tax payment in 2000.[37] There were eight jurisdictions that had two ballot initiatives. Since the first attempt failed in all eight of these jurisdictions, there is no *Priorpass* variable for Massachusetts.

Another difference in the Massachusetts analysis is that the open-space data are disaggregated to the level of the local jurisdiction. The variable *Openspace* is now defined as the local jurisdiction's percent of total land area that was undeveloped in 1999. *OSloss* is defined as the percent change in a local jurisdiction's undeveloped land that occurred between 1985 and 1999.[38]

The first column of Table 5 reports the probit model for whether a jurisdiction held a referendum. Consistent with the results for New Jersey, the coefficients on *Population*, *Popnchg*, and *Openspace* are all positive, but they are not statistically significant for Massachusetts. The effect of *OSloss*, however, is both positive and statistically significant. Jurisdictions that have lost more open space between 1985 and 1999 are more likely to have held a referenda, although the magnitude of the effect is relatively small. In particular,

[36]Beyond the observations included in *LandVote*, we obtained data on the six CPA referenda that occurred between January and July of 2004. These data are available from the Community Preservation Coalition's webpage (see footnote 35). These referenda were not included in the full data set because comparably up to date information was not available for the rest of the country.

[37]These data were obtained from the Massachusetts Department of Revenue and are available online at http://www.dls.state.ma.us/mdmstuf/proptax.htm.

[38]We obtained these data on open space, which are only available for selected years, from the Massachusetts Office of Geographic and Environmental Information. The files are available online at http://www.mass.gov/mgis.

from an average starting point of 6%, a 1% increase in open-space loss increases the likelihood of a referendum occurring by 1%.

Other statistically significant results in the probit model are the following. The negative coefficient on *Under18* indicates that a higher proportion of the population under the age of 18 decreases voter support. Similar to our New Jersey finding, higher rates of home ownership in Massachusetts also increase the likelihood of a referendum. Finally, we also find that jurisdictions with higher residential tax payments are more likely to have held a referendum. This result may reflect a general preference for the provision of local public goods.

The WLS estimates of the log-odds model are reported in the second column of Table 5. We include all of the referenda in the estimation and report clustered standard errors.[39] The surcharge rate has a negative and statistically significant effect on voter support. The magnitude of the effect is such that starting from the average of 51% of the voters voting yes, a one-percentage point increase in the surcharge rate drops the share of yes votes to 45%. This result has an important implication: given the marginal pass rate of most of the CPA referenda, the marginal effect of the surcharge rate appears pivotal to many of the election outcomes. Since the average surcharge rate is 2.4% among the CPA referenda, it appears that many unsuccessful ballot initiatives might have been successful with a more modest surcharge.

Of the three exemptions, only the low-income family and low- to moderate-income elderly exemption has a statistically significant effect. The positive sign of the coefficient is intuitive, as those who can least afford a tax increase and are likely to pay a relatively small share are those who are exempt from having to pay it. Sixty-seven percent of the referenda had this exemption. The insignificance of the other two exemptions may be due to insufficient variation in the data. Almost all of the referenda had the exemption on the first \$100,000 (84%), while very few had the commercial and industrial exemption (10%).

Both of the open-space variables are statistically significant. Voters responded more favorably to referenda in local jurisdictions where a greater percentage of the land area is open space. Nevertheless, more open-space loss within a local jurisdiction between 1985 and 1999 decreases voter support. While this surprising result is similar to our finding for New Jersey, the magnitude of the effect is very small.[40] Once again, we explore this relationship further and reestimate the WLS model with a quadratic term for *OSloss*. The coefficients and *t*-statistics on the *OSloss* variables are as follows:

$$-\underset{(2.64)}{0.041} \times OSloss + \underset{(3.06)}{0.003} \times OSloss^2.$$

Both coefficients are statistically significant. The interpretation is that voter support is decreasing in open-space loss initially, but eventually begins to increase if the open-space loss is sufficiently large. The turning point occurs at 6.8%, which is exactly the median value of *OSloss* for jurisdictions that held a referendum (the min and max values are −15.4 and 19.3, respectively). A possible explanation for this U-shaped relationship follows from recognizing that open-space loss may capture to some extent the level of economic development within a jurisdiction. Accordingly, voters may perceive low levels of open-space loss as beneficial for economic development and are thus more likely to vote against an open-space referendum. But when the open-space loss gets sufficiently large—perhaps representing economic development that begins to change the character of a community—voters respond with increasing support for an open-space referendum. While this quadratic relationship differs from the New Jersey results, which are based on county-level data rather than local-level data, the negative linear relationship is common between the two states.

Returning to the WLS model in Table 5, three other variables have statistically significant effects. The negative coefficient on *Over65* implies that a higher proportion of senior citizens decreases voter support. The positive coefficient on *Homevalue* indicates that higher home values increase the proportion of yes votes. Finally, the negative coefficient on *Avproptax* implies that higher average property tax payments decrease

[39]Once again, we only report the WLS estimates without the two-step correction for sample-selection bias. As with New Jersey, the selection bias in Massachusetts was insignificant in all of our specifications. Furthermore, we do not report estimates of the model excluding the repeat observations, as we did for New Jersey, because there are so few repeat observations in Massachusetts.

[40]One possible explanation for the pattern in both states that *OSloss* increases the likelihood of a referendum but then decreases voter support has to do with the influence of well-organized groups. Advocates of open-space protection may be well-organized and influence whether a referendum makes it on the ballot, but the preferences of these groups may differ from the greater voting population.

voter support, perhaps because voters feel their property taxes are already too high, and higher property taxes imply higher surcharges. Interestingly, this variable has the opposite sign in the probit model. This difference may be due to the fact that small coalitions within a jurisdiction can put an initiative on the ballot, but the preferences of the coalition may differ from the more general voting population.[41]

5. Conclusion

The purpose of this paper is to provide an empirical investigation of the factors that influence the appearance and success of voter referenda for open-space conservation. We take advantage of a data set that includes detailed information on all such referenda that occurred in the United States between 1998 and 2003. Combining these data with information from the U.S. Census and state-specific variables, we conduct a nationwide analysis along with focused analyses in New Jersey and Massachusetts. Six general questions motivate the paper. We reiterate these questions here to organize our main conclusions.

What factors contribute to the appearance of an open-space referendum in a jurisdiction? Across the nation, jurisdictions that have held open-space referenda differ from national averages in several respects. They tend to have greater population growth, greater household incomes, greater home values, and greater home ownership rates. Some of these results are similar to those that explain the occurrence of open-space referenda in local jurisdictions throughout New Jersey and Massachusetts. Additional results for these states relate to the affect of existing patterns of land use. Greater amounts of open space in a jurisdiction and greater amounts of open-space loss in recent years tend to increase the likelihood of an open-space referendum reaching the ballot.

How does an initiative's funding mechanism affect the way citizens vote? Voters are far more likely to vote in favor of an open-space policy that approves bond financing rather than a tax increase. Bonds are preferred to a variety of tax types, including property taxes, property tax surcharges, sales taxes, and income tax surcharges. This preference holds regardless of whether the referendum is held at the local, state, or county level. In many cases, the difference between financing with a bond or a tax determines whether a referendum passes or fails.

How responsive are favorable votes to the funding rates of an open-space initiative? Funding rates can affect a voter's incentives in two ways. Higher rates imply that each voter must pay more. At the same time, higher rates imply more open-space provision and spillin benefits for each voter. These two effects are consistent with the estimated differences in voting behavior at the local and state–county levels. At the local level—where spillin effects are likely to be small—higher funding rates tend to decrease voter support. At the state–county level—where spillin effects are likely to be large—higher funding rates tend to increase voter support.

How do socioeconomic characteristics influence preferences and therefore voting results for open-space conservation? We find evidence that collectively provided open space is a normal good. Jurisdictions with greater household income tend to be more likely to vote in favor of an open-space referendum. While property values, average property tax payments, home ownership rates, and age profiles have a significant effect in many of the econometric models, general results for these variables do not emerge across all models.

What is the effect of existing patterns of land use on voting results? In Massachusetts, more open space as a percentage of a jurisdiction's total land area tends to increase voter support. In New Jersey, where data is only available for each jurisdiction's county, the result is not statistically significant. Perhaps the most surprising result, for both states, is that more open-space loss in recent years decreases the odds of a yes vote in an open-space referendum. With nonlinear specifications, however, the results are more subtle and differ between the states. The strongest result is for Massachusetts, where the effect of the rate of open-space loss on voter support appears to be U-shaped.

What other features of a referendum affect voting results? At the national level, voter support for an open-space referendum that extends an existing policy is greater than support for a referendum that proposes a new policy. In New Jersey, having held more than one open-space referendum decreases voter support—unless one

[41]The Massachusetts initiatives can be put on a ballot one of three ways: a petition, a town meeting, or a city council decision. For all the initiatives, we collected data from the Massachusetts Community Preservation Coalition on how each referendum was put on the ballot. Including these variables in the WLS model, we found no significant effect on voting outcomes.

of the prior referenda passed, in which case voters are just as likely to support additional initiatives. In general, voters also appear more supportive of open-space policies that include provisions for local farmland preservation. Exemptions can also have a significant effect. In Massachusetts, the odds of a yes vote were greater for policies that included an exemption for low-income families and low- to moderate-income senior citizens. Finally, holding an open-space referendum outside the regular election cycle in November appears to increase voter support.

In conclusion, this paper provides new insights into the factors that influence the appearance and success of voter referenda for open-space conservation. While many of the results corroborate findings in the existing literature, other results are new. Most notably, this study provides the first investigation of how funding mechanisms and funding rates affect voter support for public acquisition of open space. We also provide evidence on the relationship between existing patterns of open space and voter support for open-space referenda. As open-space initiatives continue to gain popularity at the ballot box, the descriptive insights of this paper should prove useful for both policy-makers and advocates working in the area of land use management.

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