Explaining The Appearance and Success of Voter Referenda For Open-Space Conservation*

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October 8, 2004

* We thank Jon Bakija, Roger Bolton, Bill Fischel, and Steve Sheppard for helpful comments. We are especially grateful to Stacey Schulte for sparking the idea for this research and for assisting with preliminary data collection.

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Abstract

This paper provides an empirical investigation of the factors that influence the appearance and success of voter referenda for policies designed to promote 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, we conduct a nationwide analysis along with focused analyses of referenda that occurred in New Jersey and Massachusetts. Among the questions that we consider are the following: 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? And how do socioeconomic characteristics affect demand for public provision of open space?

JEL Classification Numbers: H41, Q38, Q58.

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 between 1998 and 2003, and approximately 80 percent of these initiatives passed.

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 demand and therefore voting results for open-space conservation? 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 Lands (TPL) and the Land Trust Alliance (LTA).[1] 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

[1] 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 excluded.

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 analyses of referenda that occurred in New Jersey and Massachusetts. Because of statewide policies that provide incentives for local jurisdictions to raise taxes for open-space conservation, there were numerous referenda in both states—237 in New Jersey and 137 in Massachusetts. For both states, we collect further Census data on all jurisdictions that did *not* hold referenda. We then estimate two models for each state: one to predict the probability that a jurisdiction has held an open-space referendum, and another to explain voting results conditional on having held a referendum.

Other researchers have investigated related questions. In a pioneering study of referenda results, Deacon and Shapiro (1975) analyze voting outcomes for a law in California to protect coastal zones from development. They find evidence that the natural coastal environment is a normal good, but the effect is not highly significant. Kahn and Matsusaka (1997) 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 (1994) 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 (2004) 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 the higher population and median household income increases the probability of a referendum. Howell-Moroney's study is a response to another paper by Romero and Liserio (2002). 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.[2]

Our paper makes four primary contributions to the literature. First, using the most comprehensive data set on open-space referenda to date, we analyze voting results using theoretically based econometric models. Second, we use variation in the financing mechanism across referenda (e.g., bonds or taxes) to investigate whether the type of mechanism proposed affects voter support for open-space acquisition. Third, we exploit variation is 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 the success of a referendum.

The results provide new insights into demand for open space and into the relationship between characteristics of an open-space policy and voter support. We find strong evidence that voters are more like to approve bonds than tax increases. Not surprisingly, funding rates also matter—with higher rates generally decreasing the odds of a yes vote. Interestingly, the opposite result emerges at the state and county levels, perhaps due to the potential for "spillin" effects. In general, we find that the factors influencing referenda outcomes differ between the state and county levels and the local level. We also find evidence that jurisdictions holding open-space referenda differ significantly from those that do not. Most notably, referenda tend to occur in wealthier and lower density communities that have experienced greater population growth. While socioeconomic and demographic variables influence where referenda occur, they have less of an effect on election outcomes. Nevertheless, we find fur-

[2] Romero and Liserio's (2002) conclusion is questionable, however, because of a number of methodological concerns. See Howell-Moroney (2004) for a detailed discussion.

ther evidence that collectively provided open space is a normal good. Other findings relate to the importance of farmland as a type of open space and to specific features of the proposed open-space policies.[3]

The remainder of the paper is organized as follows. Section 2 describes the data used in the analysis. Section 3 provides details on our econometric specification and estimation. 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 from two primary sources: the annual *LandVote* survey published by the Trust for Public Lands (TPL) and the Land Trust Alliance (LTA), and the U.S. Census online summary files for 1990 and 2000. The *LandVote* survey attempts to provide a comprehensive listing of all open-space referenda that involve the direct acquisition of undeveloped land.[4] Using the information contained in the *LandVote* survey, we generated variables for several characteristics of each open-space referendum. These variables include 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 jurisdiction, we then obtained Census data on population, population growth, population density, age profile, household income, home value, and home ownership rate.

Further data was 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. We also collected data on policy variables that are specific to

[3]Many of our results are comparable to findings in other strands of the literature on open space. These studies employ contingent valuation (Breffle, Morey, Lodder, 1998; Champ *et al.*, 2002; Vossler *et al.*, 2003), stated preferences for different types of open space (Kline and Wichelns, 1998), and revealed preferences for existing open space (Bates and Santerre, 2001). We refer to the findings of these studies where appropriate in the discussion of our results.

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

the ballot initiatives in each state. These additional variables are discussed later in Sections 4.2 and 4.3.

Of the 968 referenda in the *LandVote* survey between 1998 and 2003, a total of 857 observations were included in the final data set. 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, the information provided in *LandVote* was not sufficient to match the jurisdiction with a corresponding location in the U.S. Census (14 observations).[5] Third, the initiative was 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 is the case with a simple advisory measure (63 observations).

The 857 referenda included in the data set 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 percent 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 constitute 15 percent of the measures. The remaining referenda are divided among sales tax increases, income tax surcharges, and a category for other funding mechanisms.[6]

[5] These observations could not be clarified even after contacting TPL and receiving assistance in trying to resolve referendum locations.

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

Table 1: Descriptive statistics of open-space referenda

Panel A: Summary of referenda results by level of government

		Proportion	Proportion yes votes	
	N	Passing	Mean	Std. Dev.
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

Panel B: Summary of referenda results by funding mechanism

		Proportion	Proportion yes votes	
	N	Passing	Mean	Std. Dev.
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

Panel C: Descriptive statistics for funding rates

	Units	Mean	Std. Dev.	Min.	Max.
Proptaxrate	mills	0.282	0.370	0.001	2.5
Proptaxsurrate	percent	2.51	0.78	0.5	3
Bondrate	$10,000,000	4.66	23.50	0.0025	230
Salestaxrate	percent	0.437	0.397	0.03	0.2
Inctaxsurrate	percent	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.

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 thirty percent were in the "other" category. In contrast, 40 percent of the county-level referenda were bond issues, and 50 percent 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.[7]

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 mills 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 that 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 jurisdictions that held a referendum and compare them to the national averages. The t-statistics are based on a test of whether the sample mean is statistically different from the national average. All tests are statistically

[7] 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 degree of voter support.

Table 2: Descriptive statistics of 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: All variables are from the 2000 U.S. Census unless indicated otherwise. Standard deviations are reported in parentheses. Statistics are not reported for *Population* because the jurisdictions are not homogenous or comparable. The national average for *Density* includes only the contiguous 48 states.

significant except for the one comparing the proportions of the population over the age of 65. Compared to national averages, jurisdictions that have held an open-space referendum tend to have higher population growth and greater density.[8] They also tend to have greater household incomes, home values, and home ownership rates. In our studies of New Jersey and Massachusetts, we estimate logit models to test formally for differences between those jurisdictions that have held a referendum and those that have not.

3 Specification and Estimation

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 logit transformation of P_i, and it is often referred to as the log-odds ratio, which is commonly used in econometric models of aggregate voting results.[9]

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

$$\begin{aligned} logodds_{ist} &= \boldsymbol{\beta}_1'\boldsymbol{Mech}_i + \boldsymbol{\beta}_2'\boldsymbol{Rate}_i + \beta_3 Extend_i + \beta_4 Farm_i \\ &\quad + \boldsymbol{\beta}_5'\boldsymbol{Socio}_i + \boldsymbol{\beta}_6'\boldsymbol{Gov}_i + \delta_s + \lambda_t + \varepsilon_{ist}, \end{aligned} \tag{2}$$

where s denotes state and t denotes year; $\boldsymbol{Mech}_i$ is a categorical variable indicating the referendum's funding mechanism; $\boldsymbol{Rate}_i$ is a vector of funding rate variables that equal zero if the funding rate does not apply; $Extend_i$ is a dummy variable indicating whether

[8] 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. Later in the paper, we report results showing that *Density* is actually lower in jurisdictions that have held an open-space referendum.

[9] For examples, see Deacon and Shapiro (1975), Rubinfeld (1977), Dubin, Kiewiet, and Noussair (1992), Kline and Wichelns (1994), Kahn and Matsusaka (1997), and Vossler *et al.* (2003).

the referendum extends an existing policy; $Farm_i$ is a dummy variable indicating whether farmland is part of the proposed land acquisitions; $\boldsymbol{Socio}_i$ is a vector of 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_{ist} is a random error term.

The log-odds model specified in (2) has a microeconomic foundation. Deacon and Shapiro (1975) 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.[10] This micro foundation implies that the aggregate voting results can be used to make inferences about voters' underlying demand for open space. The validity of making such inferences has empirical support as well. A study by Fischel (1979) found little difference in a comparison between aggregate voting results and individual preferences for an environmental referendum in New Hampshire.

With the theoretical foundation of (2), the independent variables are useful for answering several questions and controlling for potentially important effects. The inclusion of $\boldsymbol{Mech}_i$ and $\boldsymbol{Rate}_i$ enables consideration of how voting outcomes respond to the proposed funding mechanism and funding rate. $Extend_i$ and $Farm_i$ will determine whether selected characteristics of the open-space proposal affect its success. $\boldsymbol{Socio}_i$ will be useful to detect factors that influence demand for open space, such as income and population density. $\boldsymbol{Gov}_i$ will indicate whether election results differ between levels of government. Finally, the state- and year-specific intercepts will control for unobserved state and year effects.

We estimate the models using weighted least squares (WLS) to account for heteroskdasticity due to the analysis of averaged, grouped data. The weight for each observation i is $(n_i\hat{P}_i(1-\hat{P}_i))^{\frac{1}{2}}$, where n_i is $Population$ and $\hat{P}_i$ is the predicted proportion of yes votes.[11]

[10] The simplification arises because, like most studies, we do not consider abstentions. Rubinfeld (1977) 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.

[11] We use ordinary least squares to obtain initial values of $\hat{P}_i$. We then iterate to convergence with WLS. See Greene (2000) for details on the use of these weights and on the method of estimation.

Using these weights implements the minimum chi-squared estimator, and the effect 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 having not held an open-space referendum, we are able to investigate the factors that influence the appearance of a referendum in a local jurisdiction. We assume that the underlying propensity of a jurisdiction to hold a referendum is given by

$$r_i^* = \alpha + \boldsymbol{\gamma}' \boldsymbol{Socio}_i + u_i. \tag{3}$$

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 $r_i^* \leq 0$. Assuming u_i has a logistic distribution, the parameters of (3) can be estimated with a logit model:

$$\Pr\left(r_i = 1\right) = \Lambda\left(\alpha + \boldsymbol{\gamma}' \boldsymbol{Socio}_i\right), \tag{4}$$

where $\Lambda(\cdot)$ is the logistic cumulative distribution function. We estimate this model for both New Jersey and Massachusetts in order to determine how socioeconomic characteristics 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

$$logodds_{it} = \phi_1 Rate_i + \boldsymbol{\phi}_2' \boldsymbol{X}_i + \boldsymbol{\phi}_3' \boldsymbol{Socio}_i + \theta_t + v_{it}, \tag{5}$$

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_{it} is a random error term. These models are estimated with WLS using the same weights as those discussed previously.

4 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 by estimating equation (2) with the nationwide sample of 857 referenda. These results are reported as the pooled model in Table 3. We also estimate equation (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.[12] 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.[13]

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.[14] A further reason for splitting the data in this manner follows from the weighted estimation, which places more weight on the

[12]We also estimated all of the models without state fixed effects; however, we do not report these results because they are very similar to those reported in Table 3.

[13]The clustering allows for arbitrary correlation among the error terms for observations within the same jurisdiction. While clustering has no effect on the coefficient estimates, it produces standard errors that are generally larger than those produced without clustering.

[14]Breffle, Morey, and Lodder (1998) and Champ *et al.* (2002) find empirical evidence that people place greater value on open space when they live closer to it.

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 Chow test comparing the separate estimates to the pooled estimates. The results indicate statistically significant differences between the local and state-county results ($F[22{,}772] = 4.49$, $p < 0.01$). 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. 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, property tax surcharges are significantly different from bonds. In the state-county model, property taxes and "other" mechanisms are 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 percent on the odds ratio.[15] This implies that, beginning from an average of 60 percent of the electorate voting yes (with an odds ratio of 1.4), the model predicts that switching from a bond to one of the taxes would decrease the proportion of yes votes to roughly 48 percent (with an odds ratio of 0.94). Note that this difference is pivotal for a referendum that requires a 50 percent majority to pass: on average, financing with a bond or a tax makes 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

[15] The percentage 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 (see Halvorsen and Palmquist, 1980).

Table 3: Nationwide, Local, and State-County WLS regressions

	(1) Pooled		(2) Local		(3) State-County	
	Coef.	Std. Err.	Coef.	Std. Err.	Coef.	Std. Err.
Proptax	-0.368***	(0.121)	-0.131	(0.199)	-0.346**	(0.135)
Proptaxsur	-0.631*	(0.386)	-0.492*	(0.331)	–	–
Salestax	-0.354*	(0.187)	-0.490***	(0.138)	-0.426*	(0.247)
Inctaxsur	-0.672**	(0.278)	-0.204	(0.263)	–	–
Other	-0.237***	(0.093)	0.065	(0.117)	-0.211*	(0.125)
Proptaxrate	0.396***	(0.155)	-0.280*	(0.159)	0.524**	(0.208)
Proptaxsurrate	-0.080	(0.091)	-0.042	(0.083)	–	–
Bondrate	0.002***	(0.001)	-0.040**	(0.017)	0.002***	(0.001)
Salestaxrate	-0.219*	(0.217)	-0.063	(0.118)	-0.093	(0.288)
Inctaxsurrate	-1.014	(0.924)	-0.882	(0.829)	–	–
Extend	0.557***	(0.122)	0.190	(0.133)	0.500***	(0.142)
Farm	-0.081	(0.059)	0.139	(0.086)	-0.095	(0.076)
Population	-0.001	(0.001)	0.038***	(0.010)	-0.000	(0.001)
Popnchg	-0.022	(0.024)	0.000	(0.012)	-0.139	(0.243)
Density	0.013	(0.022)	-0.016	(0.026)	-0.001	(0.040)
Under18	1.668	(1.449)	-1.627	(1.125)	1.679	(3.041)
Over65	1.756**	(0.816)	0.529	(0.691)	2.126	(2.005)
Income	0.078***	(0.027)	0.073***	(0.026)	0.116	(0.075)
Homeown	-1.304**	(0.571)	-0.577	(0.625)	-1.169	(1.318)
State	-0.070	(0.130)	–	–	–	–
County	-0.153**	(0.079)	–	–	-0.010	(0.187)
Constant	0.287	(0.373)	0.992***	(0.343)	-0.317	(0.803)
State dummies	Yes		Yes		Yes	
Year dummies	Yes		Yes		Yes	
Observations	857		688		169	
R-squared	0.78		0.54		0.88	

Notes: The dependent variable in all models is $logodds_i$. *Bond* is the omitted category for the funding mechanisms. *Local* is the omitted category for government level in the pooled model. All standard errors 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.

accrue in a fund with nonspecific future benefits. Third, the costs of a bond may 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.

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 decreases 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 effect of having to pay more oneself dominates; positive coefficients would suggest that the spillin effect of greater revenues dominates.

The results provide evidence in both directions. In the local model, all funding rate coefficients are negative, suggesting that higher funding rates decrease voter support at the local level. The coefficient is statistically significant for the property tax and bond rates, which are the most prevalent funding mechanisms in the data set. 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 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 evidence that voters were 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 percent of the voters voting yes, having the initiative be an extension increases the percent voting yes to 73.1 percent—a substantial increase. This result is intuitive because jurisdictions with extensions have already revealed a preference and willingness to pay for open space.

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 coefficient on *Income* is also positive and significant, implying 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 normal good (Breffle, Morey, and Lodder, 1998; Bates and Santerre, 2001), some find no significant effect (Deacon and Shapiro, 1975; Kline and Wichelns, 1994; Romero and Liserio, 2002), and one study finds that open space is generally a normal good but may become inferior at high levels of income (Kahn and Matsusaka, 1997).[16] None of the socioeconomic variables has statistically significant explanatory power in the state-county model.[17]

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 two-percent 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

[16] We attempted to replicate Kahn and Matsusaka's (1997) result by estimating all of the models with inclusion of a squared term for median household income. While the same pattern emerged, neither of the income coefficients were statistically significant.

[17] 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.

the appearance of a referendum in a jurisdiction. To accomplish this, we use socioeconomic data for all 566 local jurisdictions in New Jersey. We then estimate the logit model specified in (4) to explain whether or not a referenda occurred as a function of each jurisdiction's socioeconomic characteristics. The dependent variable equals 1 if the jurisdiction held at least one open-space referendum between 1998 and 2003, and it equals 0 otherwise.

The logit 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 because greater population growth is typically associated with more development, making open-space conservation a more salient issue in faster growing communities. The coefficient on *Density* is negative and implies that greater population density has a negative effect on the probability of a referendum occurring. This result is intuitive to the extent that greater density implies less "sprawled" development, as is often suggested.[18] This interpretation should be done with caution, however. While low population density may reflect widespread low-density development, it may also reflect high levels of undeveloped land. In the latter case, the negative coefficient on density could be the result of Tiebout sorting, whereby people who value open space highly and are therefore more likely to support ballot initiatives also settle in areas where open space already exists.

Other results from the logit 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 now turn to the WLS estimates of model (5) in order to explain election results in New Jersey. Mirroring the nationwide analysis, we include the variables *Proptaxrate*, *Extend*,

[18] Our results on population, population growth, and density are consistent with Howell-Moroney's (2004) findings in the Delaware Valley region.

Table 4: New Jersey Logit and WLS regressions

	(1) Logit		(2) WLS All Obs.		(3) WLS No Repeat	
	Coef.	Std. Err.	Coef.	Std. Err.	Coef.	Std. Err.
Proptaxrate	–	–	-0.544	(0.398)	-0.869*	(0.443)
Extend	–	–	0.001	(0.109)	-0.105	(0.156)
Farm	–	–	0.181**	(0.091)	0.278***	(0.106)
Repeat	–	–	-0.467***	(0.123)	–	–
Priorpass	–	–	0.440***	(0.166)	–	–
Population	2.450***	(0.516)	-0.224	(0.195)	-0.166	(0.176)
Popnchg	0.460*	(0.283)	0.016	(0.047)	0.116	(0.146)
Density	-0.297***	(0.066)	0.033	(0.032)	0.047	(0.034)
Under18	-1.973	(3.480)	-0.698	(1.698)	-0.728	(1.785)
Over65	0.763	(2.565)	1.442*	(0.861)	1.547	(0.967)
Income	0.472***	(0.126)	0.104	(0.075)	0.153*	(0.081)
Homevalue	-0.494**	(0.198)	-0.225	(0.142)	-0.293**	(0.141)
Homeown	0.688	(1.188)	0.039	(0.637)	0.341	(0.683)
Constant	-2.744**	(1.158)	-0.012	(0.570)	-0.522	(0.623)
Year dummies	–		Yes		Yes	
Observations	566		227		170	
Log Likelihood	-286.28		–		–	
R-squared	–		0.33		0.36	

Notes: The dependent variable in the logit model is equal to 1 if the jurisdiction ever held a referendum and 0 otherwise. The dependent variable in the WLS model is $logodds_i$. Standard errors in column (2) 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.

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 had 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 were passed. We estimate the model using all of the New Jersey referenda that were part of the Green Acres Program.[19] 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.

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 percent to 57.8 percent on average. The fact that voter support appears to be not very responsive to the property tax rate is consistent with other research that finds demand for open space to be price inelastic (Bates and Santerre, 2001). This may also explain why the effect is not statistically significant when all observations are included, in which case there is the confounding factor of repeat referenda.

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 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 ($t = 0.20$, $p = 0.84$). Thus, voting

[19] 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.

outcomes for referenda in jurisdictions that have already passed an open-space initiative are statistically indistinguishable from those having never held a referendum.

New Jersey voters are more likely to favor local farmland preservation than nonagricultural open space. The coefficient on *Farm* is positive and statistically significant, although the magnitude is small. This result corroborates Kline's and Wichelns' (1998) finding that individuals prefer farmland to most other types of open space. One possible reason is the active role that nongovernmental organizations such as the American Farmland Trust play in promoting farmland conservation.

Three of the socioeconomic variables are statistically significant. The positive coefficient on *Over65* implies that a higher proportion of senior citizens in a jurisdiction increases voter support. While this effect is not significant in the model without repeats, the effects of *Income* and *Homevalue* are statistically significant. These results follow the same pattern as that in the logit 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.

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.[20] Rather than levying a property tax millage, however, the CPA authorizes communities to levy a surcharge of up to 3 percent 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

[20] In addition to promoting open space, the CPA is intended to promote historic preservation and affordable housing. The law requires that at least 10 percent of the total funding be spent on each of the three objectives, with the remaining 70 percent allocated at the local legislature's discretion. Specific allocations were not specified prior to elections. To date, approximately 42 percent of CPA-funded projects have been for open-space and recreation (Community Preservation Coalition, 2004a).

for commercial and industrial properties.

Our analysis of local voting outcomes that were part of the CPA follows the same methodology that we use in the New Jersey analysis. We collected socioeconomic data for all 351 local jurisdictions in Massachusetts, of which 115 held a CPA referendum.[21] In order to account for the different exemptions, we collected further data from the Community Preservation Coalition (2004b) 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. 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.

The first column of Table 5 reports the logit model for whether a jurisdiction held a referendum. Many of the results mirror those for New Jersey. Referenda are more likely to occur in wealthier, larger, and faster growing jurisdictions that have lower population densities. In Massachusetts the effect of *Under18* is statistically significant, and the coefficient's negative sign indicates that a higher proportion of the population under the age of 18 decreases voter support. In contrast to New Jersey, the effect of *Homevalue* is positive and statistically significant. Differences in the financing mechanisms may contribute to this divergence. The negative coefficient on *Homevalue* in the New Jersey model was explained by the direct relationship between home value and the tax price faced by a homeowner. This relationship is less direct with the CPA's property tax surcharges, which imply tax prices that are less dependent on property values than an increase in property tax rates. As a result, *Homevalue* may be capturing less of a price effect and more of a wealth effect in Massachusetts.

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.[22] The

[21] 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 (2004b). These referenda were not included in the full data set, since comparably up-to-date information was not available for the rest of the country.

[22] We do not report estimates of the model excluding the repeat observations. Because there are so few of repeat observations in Massachusetts, the results are nearly identical to those reported in Table 5.

Table 5: Massachusetts Logit and WLS regressions

	(1) Logit		(2) WLS	
	Coef.	Std. Err.	Coef.	Std. Err.
Proptaxsurrate	–	–	-0.161***	(0.063)
Lowinc	–	–	0.260**	(0.112)
First100K	–	–	-0.117	(0.141)
Comind	–	–	-0.126	(0.125)
Repeat	–	–	-0.295	(0.189)
Population	2.124**	(0.878)	-0.061	(0.047)
Popnchg	2.583**	(1.110)	-0.018	(0.548)
Density	-0.191**	(0.096)	0.029	(0.026)
Under18	-16.049***	(5.219)	-4.841**	(2.207)
Over65	0.827	(3.672)	-5.783***	(1.472)
Income	0.346**	(0.167)	-0.086	(0.067)
Homevalue	0.509**	(0.259)	0.227**	(0.089)
Homeown	0.770	(1.760)	2.009**	(0.829)
Constant	-0.730	(1.700)	0.442	(0.593)
Year dummies	–		Yes	
Observations	359		122	
Log Likelihood	-200.91		–	
R-squared	–		0.42	

Notes: The dependent variable in the logit model is equal to 1 if the jurisdiction ever held a referendum and 0 otherwise. The dependent variable in the WLS model is $logodds_i$. Standard errors in the WLS model 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.

surcharge rate has a negative and statistically significant effect on voting outcomes. The magnitude of the effect is such that starting from the average of 51 percent of the voters voting yes, a 1 percent increase in the surcharge rate would drop the number of yes votes to 47 percent. Given the marginal pass rate of most CPA referenda, this effect appears pivotal to many of the election outcomes. Since the average surcharge rate is 2.4 percent among the CPA referenda, it appears that many unsuccessful ballot initiatives might have been successful with a more modest surcharge rate.

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 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 percent), while very few had the commercial and industrial exemption (10 percent).

While the effect of *Repeat* is not significant in the CPA model, several of the socioeconomic variables have statistically significant coefficients. The coefficients on *Under18* and *Over65* have the same sign as those in the logit model. Higher proportions of senior citizens and children not only decrease the probability of a referendum occurring; they also reduce voter support in actual elections. The effect of *Homevalue* is also similar to that in the logit model. Higher home values increase the proportion of yes votes, in addition to the probability of a referendum occurring. *Homeown* is positive and significant, perhaps because home owners are more likely to be long term residents and stand to benefit from the value of open-space being capitalized in property values.[23]

[23] Hedonic studies have found evidence that open space has a positive effect on residential property values (e.g., Irwin, 2002; Smith, Poulos, and Kim, 2002).

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, we conduct a nationwide analysis along with focused analyses in New Jersey and Massachusetts. Five 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. A similar pattern emerges in models that explain the probability of a referendum occurring in local jurisdictions in New Jersey and Massachusetts. In these models, population density is also a significant predictor of whether or not a referendum occurs: jurisdictions with lower density, which may proxy for more "urban sprawl," are more likely to have held an open-space referendum.

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 explain 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 decrease voter support. At the state-county level—where spillin effects are likely to be large—higher funding rates increase voter support.

How do socioeconomic characteristics influence demand 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 more likely to have held an open-space referendum and to exhibit greater voter support. While property values, 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 other features of a referendum affect voting outcomes? Not surprisingly, 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 of the prior referenda passed, in which case voters are just as likely to support additional initiatives. New Jersey voters are also more supportive of open-space policies that include provisions for local farmland preservation. Exemptions can also have a significant effect on voter support. 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.

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. 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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