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Is the Median Voter Decisive? Evidence of 'Ends Against the Middle' From Referenda Voting Patterns

Eric J. Brunner
Quinnipiac University

Stephen L. Ross
University of Connecticut

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341 Mansfield Road, Unit 1063
Storrs, CT 06269-1063
Phone: (860) 486-3022
Fax: (860) 486-4463
<http://www.econ.uconn.edu/>

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Abstract

This paper examines whether the voter with the median income is decisive in local spending decisions. Previous tests have relied on cross-sectional data while we make use of a pair of California referenda to estimate a first difference specification. The referenda proposed to lower the required vote share for passing local educational bonding initiatives from 67 to 50 percent and 67 to 55 percent, respectively. We find that voters rationally consider future public service decisions when deciding how to vote on voting rules, but the empirical evidence strongly suggests that an income percentile below the median is decisive for majority voting rules. This finding is consistent with high income voters with weak demand for public educational services voting with the poor against increases in public spending on education.

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1. Introduction

The median voter model has a long theoretical and empirical history within public economics. Since the pioneering work of Bergstrom and Goodman (1973), which established the conditions under which the median voter is also the voter with the median income, hundreds of studies have used the median voter framework to estimate demands for publicly provided goods and services.¹ The enduring popularity of the model stems both from its simplicity and its analytic tractability. As noted by Inman (1978), if governments act “as if” to maximize the preferences of the median income voter, the median voter hypothesis provides “a powerful starting point for predictive and normative analysis of government behavior.”

Despite the wide spread popularity of the median voter model, the key assumption that the median voter is also the voter with the median income has been repeatedly challenged.² One of the central challenges has been Tiebout sorting, whereby household sort into communities based on their demand for public services. With sorting, communities may contain both higher income households with weak preferences for public services and lower income households with strong preferences for public services. Consequently, the median preference voter may not be the voter with the median income.³ Epple and Platt (1998) develop a model that explicitly accounts for household sorting according to preferences and demonstrate that when households differ in terms of income and preferences, the median income voter is typically not pivotal. Rather, communities contain a continuum of pivotal voters differing in both income and preferences. Epple and Sieg (1999) and Epple, Romer and Sieg (2001) estimate structural models that allow for preference heterogeneity and enable them to estimate income elasticities in a model that explicitly identifies the median preference voters. Their results indicate substantial preference heterogeneity within communities, suggesting that the median preference voter is unlikely to be the voter with the median income.

¹ A review of older studies that use the median voter framework to estimate demand can be found in Inman (1979). A few of the more recent studies include, Rothstein (1992), Silva and Sonstelie (1995), Stevens and Mason (1996), de Bartolome (1997) for school spending, Schwab and Zampelli (1987) for police, Duncombe (1991) for fire, Balsdon, Brunner and Rueben (2003) for local general obligation bond issues, and Husted and Kenny (1997) for expansion of the voting franchise. The vast majority of studies are based on aggregate cross sectional data. A smaller set of studies, including Bergstrom, Rubinfeld and Shapiro (1982), Gramlich and Rubinfeld (1982) and Rubinfeld, Shapiro and Roberts (1987), use individual-level survey data to estimate demand for publicly provided goods and services.

² See Holcombe (1989) for a review of the criticisms and concerns surrounding the median voter model, and see Wildasin (1986) for an extended discussion of the assumptions required for the median voter model to be applied empirically.

³ From a statistical perspective, Tiebout sorting induces a correlation between unobserved preferences for public services and observed incomes which leads to biased parameter estimates in a Bergstrom and Goodman style demand model. This bias has been dubbed Tiebout bias by Goldstein and Pauly (1981). See Ross and Yinger (1999) for a review of the literature on Tiebout bias.

More recently, Fletcher and Kenny (2008) develop a model in which the elderly, who typically have weak preferences for local educational services, vote with the poor in support of lower levels of education spending. They demonstrate that a larger share of elderly results in a pivotal voter who is further down a community's income distribution. Similarly, Epple and Romano (1996a, 1996b) demonstrate that when there exist private alternatives to public goods or when public goods can be supplemented with private purchases, an equilibrium exists where the pivotal voter has an income that lies below the median. All three papers describe situations where low- and high-income voters with weak preferences or demand for public service provision form a coalition to oppose the preferences of middle and high-income voters with strong preferences; situations Epple and Romano (1996a, 1996b) describe as "ends-against-the-middle."

In light of these challenges, numerous studies have attempted to test whether the voter with the median income is empirically relevant for describing local public service provision. Pommerehne and Frey (1976), Pommerehne (1978), Inman (1978), Turnbull and Djoundourian (1994) and Turnbull and Mitias (1998) evaluate the performance of the median voter model by examining whether the use of median income in local public service demand regressions outperforms other specifications (such as replacing median income with mean income). The results of those studies generally support the hypothesis that the median income voter is decisive.⁴ On the other hand, Aronsson and Wikström (1996) test the predictive power of a model where the median income voter is assumed to be decisive against a more general statistical alternative. Their results lead them to reject the hypothesis that the voter with the median income is decisive.

A common feature of all these prior tests is that they rely on aggregate cross-sectional data to identify a relationship between public service expenditure levels and some measure of community income. These studies are likely biased because communities differ across a variety of dimensions including unobserved preferences for public services, the cost of providing public services, etc; and these differences are likely correlated with the distribution of income in each community.⁵

In this paper, we propose an entirely new approach for testing the median voter hypothesis. We examine vote returns from a unique pair of California referenda that proposed changing the rules under which public spending decisions are determined. The first referendum, which failed, proposed to lower the required vote share for passing local educational bonding initiatives from 67 to 50 percent, and the second referendum, which was held only eight months later and passed, proposed lowering the vote

⁴ Using a revealed preference approach, Turnbull and Chang (1998) also find that local governments act "as if" to maximize the utility of the median income voter.

⁵ For example, Schwab and Zampelli (1987) find that studies of public service demand that fail to take into account the impact of community characteristics on the cost of public service provision can yield very misleading results. See Ross and Yinger (1999) for a survey of studies that document cost heterogeneity across jurisdictions, as well as recent additional studies by Duncombe and Yinger (2005) and Reschovsky and Imazeki (2003).

requirement from 67 to 55 percent. Thus, assuming demand is monotonically increasing in income, the first referenda would have changed the identity of the decisive voter from the voter in the 33rd percentile of the income distribution to the 50th percentile while the second referenda would have changed the identity of the decisive voter from the voter in the 33rd percentile to the voter in the 45th percentile. Using the results from these two referenda, we test whether people vote “as if” future spending decisions will be based upon the preferences of the newly proposed decisive voter by examining whether the change in the fraction of ‘yes’ votes cast in the two elections can be explained by the implied change in the newly proposed decisive voter’s income, i.e. the difference between the 50th and 45th percentile income in a jurisdiction.

Unlike previous tests of the median voter hypothesis, where public service spending is used to infer a relationship between the median voter’s preferences and outcomes of the political process, our test infers that a median voter relationship holds because voters act as if the relationship holds when they cast their ballots to determine voting rules for choosing the level of public services provided. Consequently, our test avoids the fundamental problem of measuring the actual services demanded by voters within a jurisdiction which may be poorly proxied by the measures used in previous studies, such as expenditures per capita.⁶ Furthermore, by regressing changes in the fraction of ‘yes’ votes between the referenda on changes in the income associated with the decisive voter in each district, we are able to difference out school district unobservables that are likely correlated with the distribution of income within a district and likely bias prior cross-sectional tests of the median voter hypothesis.

We find a strong relationship between the income distribution of a school district and the change in the fraction of ‘yes’ votes between the two referenda. This relationship, however, appears to arise most strongly from the influence of the income difference between 40th and 35th percentile on voting rather than the 50th to 45th percentile. Specifically, while the income difference between the 50th to 45th percentile can explain changes in voting between the two referenda, when we run a “horse race” between the changes in income between the 50th to 45th percentile and the 40th and 35th percentile, the difference between lower percentiles entirely captures the systematic relationship between the income distribution and voting.⁷ These findings persist across a series of specifications controlling for changes in turnout and political representation between the two referenda, differences between small and large school districts, and demographic differences between school districts. The estimated relationship also persists for

⁶ For example, as noted by Behrman and Craig (1987), “people pay taxes based on the city-wide amount of purchased inputs, but base their demand and voting behavior on the perceived level of neighborhood service output.” Thus, to the extent that the services produced differ substantially across jurisdictions given the same public inputs, public spending will provide a poor proxy for public service provision.

⁷ Note that the theory of referenda voting on which our empirical model is built holds as long as referenda voters’ anticipation of future public service levels can be characterized by the demand of voters at a specific income percentile. This percentile need not be the 50th percentile.

constant income elasticity models that allow for heterogeneity in the distribution of preferences across school districts. Furthermore, we find that the relationship between changes in the decisive voter's income and changes in vote shares does not hold for two counterfactuals estimated by replacing school districts with alternative definitions of jurisdiction based on census tracts and state assembly districts.

Having rejected the hypothesis that the median income voter is decisive, we proceed to examine whether our results are consistent with preference heterogeneity leading to an "ends-against-the-middle" outcome similar to the type described by Epple and Romano (1996a, 1996b) and Fletcher and Kenny (2008). We split our sample based on the fraction of individuals in a district that are high income and yet are expected to have weak preferences (i.e. households without children) or low demand (i.e. households with children in private school) for public education services and find evidence that points towards a *lower* income percentile decisive voter (further from the median) for districts with a greater fraction of high-income/low-demand households and a *higher* income percentile decisive voter for districts with a smaller fraction of such households.

In light of these findings, we estimate a final model in the spirit of Epple and Platt (1998) and Epple and Sieg (1999) where we explicitly recognize that each community contains both high and low demand voters. We use the share of voters who are past traditional childbearing age (45 years or older) and do not have school-age children to represent the fraction of low demand voters in a district.⁸ We then estimate our change in fraction of "yes" votes simultaneously with two additional equations that determine the decisive voter's demand for spending under the two voting rules (55 and 50 percent thresholds). These two additional equations allow us to identify the decisive voter income percentile for the low and high demand households under each voting rule. The model performs as expected with increases in the share of low demand voters being associated with lower decisive voter income percentiles for the high demand group. The model implies that low demand individuals have demands for public services that are 45% lower than other voters. The model also yields estimated income elasticities of between 0.695 and 0.876 which are very comparable to the estimated elasticities between 0.615 and 0.878 in our constant elasticity of demand model using the lower income percentiles, but substantially below elasticity estimates based on the jurisdiction median income.

Finally, based on our structural constant income elasticity models, we find evidence that the relationship between the decisive voter's demand and support for voting rules is weaker in more heterogeneous districts; evidence that provides empirical support for the referenda voting models

⁸ Share of adults with children in private school might provide an alternative proxy, but this variable is much less empirically relevant. Only about 11% of households with school aged children have children in private school, and that implies that only about 3% of households overall have children in private school. Further, families with children in private school are spread widely across the income distribution so that when considering high income households with children in private school, who would presumably vote for high levels of public services without the private school option, the relevant share of households is well below 2%.

developed by Romer, Rosenthal and Munley (1992) and Rothstein (1994). Empirical support for referenda models of this type is especially important given that these models predict less electoral support for spending initiatives in the presence of greater voter heterogeneity; a prediction that is consistent with empirical findings of Alesina, Baqir and Easterly (1999).

2. Conceptual Framework

Prior to 2000, local school bond measures in California required a two-thirds supermajority to pass. If voters approved a bond issue, the bonds were then repaid with local property tax increases that remained in effect until the bonds were fully repaid. In 2000 Californians voted on two statewide initiatives designed to ease this supermajority vote requirement. In March of 2000 Californians voted on Proposition 26, an initiative that would have reduced the vote requirement on school bond measures to a simple majority. The proposition garnered the support of only 48.7 percent of voters and thus failed. In November of 2000 Californians voted on Proposition 39, an initiative that was nearly identical to Proposition 26 except it called for reducing the vote requirement on local school bond measures to 55%. This time California voters approved the measure with 53.3 percent of voters supporting Proposition 39.

To motivate our empirical work, we begin by examining the implications of the median voter model for the behavior of voters in a referendum on voting rules. Specifically, we develop a simple voting model based on Romer, Rosenthal and Munley (1992) and Rothstein (1994) in order to illustrate the relationship between support for a change in required vote share and the income of the decisive voter. Let S_{ij}^* denote the desired level of school spending of individual i located in school district j . The individual votes in favor of a decrease in the vote share required to pass spending referenda to P if and only if S_j^P , the spending level under the new vote share, is preferred to S_j^0 , the spending level under the current, higher vote share requirement.

Following Rothstein (1994), we parameterize individual preferences for school spending using the desired spending level S_{ij}^* so that an individual's indirect utility function $V(S_j | S_{ij}^*)$ is maximized when district spending level $S_j = S_{ij}^*$. We allow districts to vary in terms of resident's preference for school spending by assuming that the distribution of preferences in district j is distributed around a district mean preference S_j^* or

$$S_{ij}^* = S_j^* + \mu_{ij}, \quad (1)$$

where μ_{ij} is a random disturbance. Assuming preferences over public service levels are single peaked, a unique α_j^P exists so that $V(S_j^0 | \alpha_j^P) = V(S_j^P | \alpha_j^P)$ where V is a voter's indirect utility function conditional on their preferences for the public service.⁹ A voter supports the referendum to lower the vote share requirement (presumably increasing public service levels) if and only if $S_{ij}^* > \alpha_j^P$.

The function α_j^P only varies across communities based on the values of S_j^0 and S_j^P because all households in the economy are assumed to have common preferences except for the preference shifter S_{ij}^* .

Using a linear parameterization, α_j^P may be written as

$$\alpha_j^P = \alpha(S_j^0, S_j^P) + \nu_j = (1 - \delta)S_j^0 + \delta S_j^P + \nu_j, \quad (2)$$

where δ is between zero and one. Equations (1) and (2) imply that a voter supports the reduction in the vote share requirement if and only if

$$S_{ij}^* - [(1 - \delta)S_j^0 + \delta S_j^P] > \nu_j - \mu_{ij}. \quad (3)$$

If ν_j and μ_{ij} follow independent type 1 extreme value distributions, then the difference between the unobservables in equation (3) has a logistic distribution with a cumulative distribution function of

$$F(x) = 1 - \exp \left[- \exp \left(\frac{x - (S_{ij}^* - [(1 - \delta)S_j^0 + \delta S_j^P])}{\beta} \right) \right], \quad (4)$$

where the variance of the distribution is $\frac{1}{3}\beta^2\pi^2$ and equals the sum of the variances of ν_j and μ_{ij} ,

assuming that the two disturbances are independent (Johnson, Kotz, and Balakrishnan, 1995).

Let p_{yes_j} denote the fraction of voters in district j that prefer S_j^P to S_j^0 . If we assume that β is constant across communities and without loss of generality initialized to one, the log-odds ratio for referendum k can be expressed as:

⁹ Specifically, we assume that indifference curves are convex over public service levels so that given a well-behaved community budget constraint, $V(S_j | S_{ij}^*)$ is a concave function of S_j . See Rothstein (1994), Epple and Romano (1996a, 1996b), Balsa and Brunner (2005), as well as many other earlier papers that impose such assumptions.

$$\ln\left(\frac{pyes_j^k}{1-pyes_j^k}\right) = c_{0j} + c_1 S_j^* - c_2 S_j^0 - c_3 S_j^P + \varepsilon_{jk}, \quad (5)$$

where c_{0j} captures idiosyncratic referenda invariant district attributes that influence voting, c_1 , c_2 , and c_3 are all non-negative as found by Rothstein (1994) and ε_{jk} represents referenda specific district unobservables. Equation (5) suggests a simple differencing estimation strategy. Specifically, if the first referendum imposed a 50% vote share for local spending measures and the second referendum imposed a 55% vote share, the difference in the log-odds of the fraction of voters that support the two referenda in district j is:

$$\ln\left(\frac{pyes_j^2}{1-pyes_j^2}\right) - \ln\left(\frac{pyes_j^1}{1-pyes_j^1}\right) = -c_3 (S_j^{55} - S_j^{50}) + (\varepsilon_{j2} - \varepsilon_{j1}), \quad (6)$$

where both the unique district mean preference for public service levels, the default level of public service provision, and referenda invariant district attributes drop out of the model.

The assumption that β is constant across communities, however, is quite strong given other assumptions in the model. Specifically, once μ_{ij} is restricted to follow an independent type 1 extreme value distribution, σ_μ^2 , the variance of μ_{ij} , must be positively related to the difference between the 50th and 55th percentile demands since only an increase in the variance can create additional spread at the center of the distribution holding the form of the distribution fixed. Therefore, the assumption that β is constant requires that σ_v^2 , the variance of v_j , fall by the exact same amount as any increase in σ_μ^2 . Further, there are reasons to believe that v_j and μ_{ij} are correlated across communities and that σ_v^2 might also depend upon factors that influence the heterogeneity of preferences.¹⁰ To allow for heterogeneity in the distribution of preferences across districts we therefore extend equation (6) as follows:

$$\ln\left(\frac{pyes_j^2}{1-pyes_j^2}\right) - \ln\left(\frac{pyes_j^1}{1-pyes_j^1}\right) = -\frac{c_3 (S_j^{55} - S_j^{50})}{\beta_j} + (\varepsilon_{j2} - \varepsilon_{j1}), \quad (7)$$

¹⁰ v_j captures community specific errors describing the indifference point between the proposed and initial spending levels, S^P and S^0 , which obviously depends on S^* . Different values of S^P and S^0 fall in different regions of the indirect utility function with different curvature implying they may still influence the distribution of v_j . Naturally, S^P and S^0 depend upon the same unobservables that influence the distribution of μ_{ij} .

where β_j describes the unique standard deviation of the preference distribution for district j .

As equation (7) reveals, inter-district differences in preference heterogeneity scales the influence that the decisive voter's demand has on referenda vote outcomes. Specifically, the influence of the decisive voter's demand on the determination of voting rules is smaller in more heterogeneous districts (i.e. districts with a larger β). This result mirrors the results of Romer, Rosenthal and Munley (1992) and Rothstein (1994) who demonstrated that, all else equal, greater preference heterogeneity within jurisdictions results in lower approval margins for local referenda.¹¹ In our model, this result can be observed from equation (4) where the impact of support for a proposed level of public services S^P on the likelihood of voting yes is scaled down by preference heterogeneity as captured by β . In the present context, equation (7) suggests that differences in the desired level of spending of the 50th and 55th percentile voters should have a smaller impact on the change in the fraction of 'yes' votes in districts with significant preference heterogeneity and a greater impact in districts with relatively little preference heterogeneity.

3. Empirical Methodology

In order to operationalize equation (6), maintaining the assumption that preferences are singled peaked and at least initially maintaining the assumption that β is constant across districts, we assume for referendum k that the implied future level of public service is the public service demand of a voter with the income implied by the income percentile associated with the vote share required under the referendum. For a 50 percent vote share, the standard median voter result applies so that the demand of the median preference voter is a Condorcet winner (Black, 1948a). Therefore, for the first referendaum, we allow the future level of public service to be represented by the public service demand of the median income voter in the school district.

For a 55 percent vote share, which is the super-majority requirement associated with the second referendum, a set of feasible δ -majority winners arises that can both defeat the original status quo¹² by the required 55 percent vote share and cannot be defeated by any other service level in pairwise competition. The set of feasible winners is in the interval between the public service demand of the 45th percentile voter and a higher service level which leaves the 45th percentile voter indifferent between that level and the status quo (Black, 1948b). Gradstein (1999) demonstrates that among these feasible winners the

¹¹ See Proposition 2 in Romer, Rosenthal and Munley (1992).

¹² We assume that the pre-referenda service level is below the desired service level of 45th percentile voter, which is consistent with the observation that passage of the referenda led to higher spending, which would not have occurred if the status quo was above this level since then the status quo would have been a δ -majority winner.

demand of the $(1 - \delta)$ percentile preference voter will be selected for a super majority requirement of δ if the government starts by proposing the status quo followed continually by ballots associated with higher levels of spending until a referenda fails to achieve the required super-majority. Under these voting rules, all voters will rationally choose to vote sincerely and the referenda will fail for services levels just passed the level demanded by the $(1 - \delta)$ percentile preference voter.¹³ For a 55 percent vote share, we therefore follow Gradstein (1999) and allow the future level of public service to be represented by the public service demand of the voter with the 45th percentile income in the school district.

Assuming a linear form for public service demand yields

$$S_j^{P_k} = S(y_j^{100-P_k}) = b_1 + b_y y_j^{100-P_k}, \quad (8)$$

where P_k is the required vote share for referendum k , $y_j^{100-P_k}$ or y_{jk} for short is the income at the decisive percentile, and b_y is the parameter describing the responsiveness of demand to income.

Substituting equation (8) into equation (6) and rearranging yields

$$\ln\left(\frac{pyes_j^2}{1 - pyes_j^2}\right) - \ln\left(\frac{pyes_j^1}{1 - pyes_j^1}\right) = (d_1 - d_2) + d_y(y_{j1} - y_{j2}) + (\varepsilon_{j1} - \varepsilon_{j2}), \quad (9)$$

where the d_y parameter is just the b_y parameter from equation (8) multiplied by the negative term $-c_3$ and the difference between d_2 and d_1 allows the mean of the district unobservable, ε_{jk} , to vary across referenda. The median or decisive voter model predicts that d_y should be positive since public service demand increases with income. Specifically, each referendum is assumed to be supported by all voters who prefer the new higher level of spending to current spending based on the two-thirds vote requirement. A larger income difference between the 50th and 55th percentile voter (i.e. the 50th and 45th percentile income), implies a larger reduction in new education spending or a smaller increase in spending over current levels, which is then supported by more voters.

The first difference specification in equation (9) eliminates unobserved differences across districts in the average preference for public services, political leaning, time invariant differences in

¹³ Similar, but more realistic voting rules, also yield the 45th percentile voter as decisive. Specifically, if government chooses a high level of public services for the initial referendum and continuously reduces the level in subsequent ballots until a proposed service level obtains the required super-majority, then all voters with preferences above the proposed service level vote sincerely, but voters with preferences below vote strategically knowing that even though they might prefer the proposed level to the status quo referenda failure will result in a marginal lowering of the public service level. Again, voting concludes when the proposed service level just reaches the demand of the 45th percentile preference voter.

turnout rates, as well as a host of other idiosyncratic differences that affect voting and additively enter the estimation equation. Consequently, equation (9) eliminates time invariant factors that influence voting patterns and might be correlated with the income distribution and thus bias a cross-sectional test of the median voter model. Nevertheless, our first difference specification does not address the concern that changes in the decisive voter's income between referenda may be correlated with unobservables that affect the change in vote share between referenda. In order to control for such factors, additional models of change in vote share are estimated including linear controls for voter turnout and other attributes intended to capture changes in the composition of voters between the two referenda, such as district size and the fraction of residents that are college educated.

As discussed earlier, the median income voter may not be decisive under a majority rule and accordingly we consider the income difference between other income percentiles. Specifically, we run regressions that include the income difference between the 50th and 45th percentiles along with an additional income difference, such as the difference between the 45th and 40th or the 60th and 55th percentiles. The winner of these so called "Horse Races" provides evidence concerning the income of the decisive voter under majority voting.

Heterogeneous Preference Distributions with Constant Elasticity Demand

By specifying public service demand to be a simple linear function of the decisive voter's income we are able to derive an estimation equation that is linear in the parameters, which is attractive from an estimation standpoint because unobservables in voting behavior and preferences are differenced away. However, the linear model assumes that preference heterogeneity, as captured by β in equation (7), is constant across school districts. In order to relax this assumption, we allow the distribution of preferences across jurisdictions to vary as a parametric function of a vector jurisdiction attributes, Z_j , that capture the heterogeneity in preferences within school district j . Moreover, the common practice in the literature is to assume public service demand is characterized by constant income elasticity which implies

$$S_j^{P_k} = b_2 \left(y_j^{100-P_k} \right)^\theta, \quad (10)$$

where θ is the income elasticity of public service demand. Substituting equation (10) into equation (7) and rearranging yields

$$\ln \left(\frac{pyes_j^2}{1 - pyes_j^2} \right) - \ln \left(\frac{pyes_j^1}{1 - pyes_j^1} \right) = d_3 \exp[\phi Z_j] [y_{j1}^\theta - y_{j2}^\theta] + (\varepsilon_{j1} - \varepsilon_{j2}), \quad (11)$$

where the d_3 parameter is the b_2 parameter from equation (10) multiplied by the term $-c_3$ from equation (7), and ϕ is a vector of parameters on the heterogeneity variables. We expect the income elasticity θ to be positive while we expect ϕ to be negative since increased preference heterogeneity reduces the influence of the decisive voter's demand on vote outcomes. Because equation (11) is nonlinear in the parameters, we use non-linear least squares to estimate the parameters of this model. Furthermore, as with the linear model, we consider additional specifications that include a host of linear controls that explain changes in vote share.

4. Data

We obtained data on vote outcomes for Propositions 26 and 39 from the Statewide Database, maintained by the Institute of Governmental Studies at the University of California, Berkeley. The database contains aggregate vote outcomes and voter registration information for all statewide primary and general elections held in California since 1990. The primary unit of analysis in the statewide database is the Census block. We aggregated the block level vote tallies and voter registration information up to the school district level.

In the empirical framework developed in Section 3, the difference in vote shares between Propositions 39 and 26 is a function of the difference between the 50th and 45th percentile income in a school district. To construct estimates of the income percentiles, we used district-level data from the 2000 Census on the distribution of household income. Specifically, the 2000 Census contains information on household income grouped into 17 income categories. We used this grouped income data and linear interpolation to estimate the 50th and 45th percentile level of income in each district. Using similar methods, we develop measures of the income difference for other percentiles, such as the difference between the 40th and 35th or 60th and 55th income percentiles.¹⁴

We also include a number of additional variables in several of our empirical specifications. The first variable is the difference in voter turnout between Propositions 26 and 39. Following Coate, Conlin and Moro (2008) and Coate and Conlin (2004) among others, we define voter turnout as the fraction of eligible voters (i.e., voting age population) in each district that voted on Proposition 39 and Proposition 26 respectively. We include the difference in voter turnout to account for the potential impact changes in

¹⁴ We have used a similar approach to construct measures of the 50th and 45th percentile tax prices. Unfortunately, district-level data on the assessed value of owner-occupied homes in California is unavailable, and we are forced to calculate our tax price variables using data from the 2000 census on the distribution of house values, which is undesirable because in California home values are likely to vary dramatically from assessed values due to Proposition 13, which prohibits the reassessment of homes for property tax purposes except when the house is sold. Nonetheless, we have estimated many of our models including controls for tax price, as well as income. While noisy, our tax price estimates are quite reasonable, and all other findings are robust to the inclusion of controls for tax price.

voter turnout may have on vote outcomes between the two elections. The second variable is the difference in the fraction of registered Republicans between Propositions 26 and 39 and the third variable is the difference in the fraction of registered Democrats between the two propositions. We include these two variables to control for systematic changes in the ideological composition of voters between elections.

In addition to the difference control variables described above, we also include several level control variables in some of our empirical specifications. The first three variables are district size fixed effects. Specifically, we sorted districts into four equally sized groups based on total population, and created three indicator variables that take the value of unity if a district is in the second, third or fourth quartile of district size respectively. The next five variables describe the demographic composition of a school district. We include controls for (1) the fraction of the population with a bachelor's degree or higher, (2) the fraction of homeowners, (3) the fraction of households that are White and non-Hispanic, (4) the fraction of households with children, and (5) the fraction of the voting age population age 65 or older. These variables were selected to capture factors that have been found in earlier literature to influence voting and public spending decisions.

To implement the specification given by equation (11), which allows the distribution of preferences to differ across districts, we also developed three variables designed to measure preference heterogeneity within districts. As discussed in section 3, district preference heterogeneity, as captured by β , is likely related to the distribution of income. Consequently, we include a Gini index of income inequality. Preference heterogeneity is also likely to be related to the fraction of households who do not have school-age children and/or do not plan on having children in the future; households that we shall refer to as low-demand households. We use the fraction of voters who are age 45 or older and do not have school-aged children as our measure of the share of low-demand households and include the variance of this measure as an additional control for preference heterogeneity in a district.¹⁵ Finally, preference heterogeneity is also likely to be related to the degree of racial heterogeneity (Alesina, Baqir and Hoxby, 2004). Following Urquiola (2005), the racial heterogeneity index we employ is:

$$I_j^{race} = 1 - \sum_{r=1}^R R_{rj}^2, \text{ where } R_{rj} \text{ is racial group } r\text{'s share of the population in school district } j. \text{ Greater}$$

values of this index are associated with greater racial heterogeneity.

¹⁵ Data on the fraction of voters age 45 or older was obtained from the Statewide Database, which includes information on the age of registered voters in each voting precinct. To construct the fraction of voters with no school age children, we multiplied the number of voters age 45 or older by the fraction of the population age 45 or older with no school-age children. We then divided this number by the total number of registered voters in each district.

Our data have a number of limitations. The first limitation concerns school districts with overlapping boundaries. Specifically, California contains three types of school districts: unified districts, elementary districts and high school districts. The boundaries of the latter two types of districts overlap: one high school district typically contains two or more elementary districts. Thus, in non-unified districts there are really two decisive voters, the decisive voter for the elementary school district and the decisive voter for the high school district into which the elementary district feeds. Consequently, in non-unified districts it is unclear how one should measure the income of the proposed decisive voter. We therefore restrict our sample to unified school districts. The second limitation concerns missing data. Data on the fraction of voters supporting Proposition 26 is unavailable for 17 of the 323 unified school districts operating in California in 1999-2000.¹⁶ We exclude these 17 districts from our analysis leaving a final sample of 306 unified school districts.

Table 1 provides means and standard deviations for the variables used in the analysis. For variables that enter our model as differences, the summary statistics are reported separately for Propositions 26 and 39 respectively. As expected, the increase in the vote requirement from 50 to 55 percent is associated with a greater percentage of voters supporting the referendum and lower decisive voter income. The change from a March election (Prop 26) to a November election (Prop 39) also increases turnout from 28 to 43 percent of eligible voters.

{Insert Table 1 Here}

5. Results

Regression results for the change in vote share using the linear demand specification in equation (9) are presented in Table 2. The first column presents the basic model that controls only for the change in the decisive voter's income. The second and third columns contain results from models that include controls for changes between the two elections in turnout and party affiliation among registered voters and those controls plus jurisdiction size fixed effects, while the fourth column contains results based on a model that includes the controls used in column three plus additional controls for district demographic attributes. As expected, all four regressions imply a strong positive relationship between the change in the decisive voter's income and the change in the fraction of 'yes' votes between Propositions 39 and

¹⁶ 15 of the 17 districts with missing vote data were located in the counties of Monterey, Humboldt and San Luis Obispo which did not report vote tallies to the Statewide Database for Proposition 26. The remaining two districts are small rural districts that had substantial missing observations on voting for Proposition 26.

26.¹⁷ The estimated coefficients on the change in the decisive voter's income range from 0.143 to 0.232 and are all statistically significant at the five percent level. These point estimates suggest that a large fraction of the change in vote shares between the two propositions can be explained by the change in the decisive voter's income. Specifically, based on equation (9) our model predicts that the implied change in the decisive voter's income is consistent with a 2.3 percentage point increase in the percent voting yes in model 1, and a 2.8, 1.7, and 2.5 percentage point increase in the percent voting yes in models 2, 3, and 4 respectively.¹⁸ Given that the actual increase in percent voting yes was 4.3 percentage points, our model predicts that between 39% and 65% of the change in vote shares between the two propositions can be explained by the change in the decisive voter's income.¹⁹

{ Insert Table 2 Here }

The top panel of Table 3 presents the results of our “Horse Races” between the median income voter for a majority voting rule and alternative income percentiles. The first two columns present coefficient estimates and standard errors for the change in income between the 50th and 45th percentile and the change in income between two other percentiles that are separated by 5 percentage points. The third and fourth columns report the p-values associated with non-nested hypothesis tests based on the null hypothesis that the “correct” model is either the one based on the 50th and 45th income percentile difference or the other income percentile difference listed in each row of the table. These non-nested hypothesis tests are constructed using the *P* test developed by Davidson and MacKinnon (1981, 1982). Note that the income differences between the 40th to 35th percentiles and the 35th to 30th percentiles clearly dominate the income difference between the 50th and 45th percentiles. In the “Horse Race” regressions the estimated coefficients on the income differences between the 40th to 35th and the 35th to 30th percentiles are relatively large in magnitude and statistically significant while the estimated coefficients on the income difference between the 50th and 45th percentile are small in magnitude and statistically insignificant. Similarly, in columns 3 and 4, we reject the null hypothesis that the correct model is the one based on the income difference between the 50th to 45th percentiles and fail to reject the null hypothesis that the correct model is the one based on either the income difference between the 40th to 35th

¹⁷ We also divided the sample into three subsamples based on the size of the increase in turnout between the two referenda. The estimated effects were similar in magnitude, and we could not reject the hypothesis that the effect of income was the same across these three subsamples. Further, we developed school district controls based on the correlation between change in turnout and neighborhood demographics at the block group level as a proxy for the change in voter demographics between referenda. The inclusion of these controls has no impact on our estimates.

¹⁸ The predicted change is estimated by calculating the change in log-odds for each school district in the sample and translating this change into a predicted change in share voting yes based on the actual share voting yes for proposition 26 in the school district.

¹⁹ The actual increase in the percent voting yes of 4.3 percentage points is based on the sample of unified school districts and thus differs from the statewide increase in the percent voting yes which was 4.5 percentage points.

or 35th to 30th percentiles. Thus, our results provide evidence against the hypothesis that the median income voter is decisive in majority rule referenda over local public service provision.

The bottom panel of Table 3 presents the results of “Horse Races” and various non-nested hypothesis tests that compare the 40th to 35th percentile income difference to other percentile income differences.²⁰ The income difference between 40th to 35th percentile clearly dominates all other percentile income differences except for the 35th to 30th, and as in the top panel, the effect size for 40th to 35th is a little bigger than the effect size for 35th to 30th. We interpret these results as implying that our data are consistent with a decisive voter near the 40th income percentile for a majority voting rule. When we re-estimate the four specifications presented in Table 2 using the change in income between the 40th and 35th percentiles rather than the 50th and 45th percentiles, we obtain coefficient estimates on the change in the decisive voter’s income that range between 1.89 and 2.87 (see Table 4 below). These point estimates imply that the change in the decisive voter’s income is consistent with between a 2.1 and a 3.2 percentage point increase in the percent voting yes between the two referenda. Thus, based on the difference between the 40th and 35th percentile incomes, our model predicts that between 49% and 74% of the change in vote shares between the two propositions can be explained by the change in the decisive voter’s income.

{ Insert Table 3 Here }

Counterfactual Analysis

In order to further test whether we have truly identified a causal relationship between changes in the decisive voter’s income and changes in vote shares, we conduct two counterfactuals. The logic behind our counterfactuals is simple: if the relationship we have identified is truly causal, then it should hold for school districts (which would have been directly affected by the outcomes of Propositions 26 and 39) but it should not hold for other political or geographic entities. For example, while we expect the income difference between the 40th and 35th percentile voter in a *school district* to explain differences in vote shares within school districts we would not expect the income difference between the 40th and 35th percentile voter in a *census tract* or a *state assembly district* to explain differences in vote shares within those geographic/political entities. That is, for political/geographic entities other than school districts, the income difference between the 40th and 35th percentile voter should be uncorrelated with changes in vote shares.

²⁰ We focus on the 40th to 35th percentile change (rather than the 35th to 30th) since the coefficient estimate on the income difference between the 40th and 35th percentiles reported in the top panel of Table 3 is slightly larger than the coefficient estimate for the 35th to 30th percentile. Results using 35th to 30th are similar.

Our rationale for choosing census tracts and State Assembly Districts (SAD) is based on their size and their lack of relevance for the provision of any local public services. Census tracts tend to be much smaller than school districts while state assembly districts tend to be larger than school districts. While some school districts such as Los Angeles Unified contain several SAD's, California contains a total of 80 SAD's relative to approximately 1,000 school districts. Thus, our counterfactuals cover geographic/political entities that are both smaller and larger than school districts on average. Further, since neither of these geographic regions represents a level of local government, the decisive voter income variables should not be related to any unexpected fiscal implications of Propositions 26 and 39.

To implement our counterfactuals we estimate models identical to those reported in Table 2, except that we use the 40th and 35th income percentiles and calculate those percentiles for either census tracts or state assembly districts. For example, our counterfactual involving census tracts utilizes information for 6,891 census tracts on vote shares in a census tract, income differences between the 40th and 35th percentile voter in a census tract, etc. Similarly, our counterfactual involving state assembly districts utilizes information for the 80 SAD's in California on vote shares within SAD's, income percentiles within SAD's, etc.²¹ In addition, we also present estimates for census tract models that include district fixed effects. District fixed effects are included to insulate the estimates against the systematic across district variation that drives the estimates in the school district sample.²² Naturally, the school district fixed effects could also contaminate our SAD estimates (in fact some mid-sized school districts essentially are SAD's), but a natural analog to the census tract fixed effects model does not exist because some SAD's contain many school districts while others are entirely contained within school districts.

Results for the counterfactuals are reported in Table 4. In the interest of brevity, we report only the estimated coefficients on the income difference variable. The first column of Table 4 replicates the school district results reported in Table 2 except the income difference is based on the decisive voter percentile identified in Table 3. The second, third and fourth columns present our counterfactuals based on state assembly districts, census tracts and a census tract model with district fixed effects, respectively. The four panels presented in Table 4 correspond to the four models listed in Table 2. The results reported in Table 4 are quite striking. In all our counterfactuals the estimated coefficients on the difference between the 40th and 35th percentile income are significantly smaller than the estimates for school districts with the exception of one model for the State Assembly Districts, where the estimate is very noisy. Furthermore, all the estimated coefficients on the difference between the 40th and 35th percentile income

²¹ Similar results arise using the difference between the 50th and 45th percentile incomes.

²² Standard errors for this model are also clustered at the school district level because heteroscedasticity can bias the estimation of standard errors in fixed effect models.

in our counterfactuals are statistically insignificant. Thus, the results reported in Table 4 give us increased confidence that our results are capturing a relationship between changes in the proposed decisive voter's income and voting patterns that is unique to school districts. They also provide strong evidence that our results are not being driven by some other unobserved factor, such as differences in voter turnout or unobservable community characteristics that might persist at a variety of levels of spatial aggregation.²³

{Insert Table 4 Here}

Heterogeneous Preference Distributions with Constant Elasticity of Demand

Nonlinear least squares estimates for the constant elasticity of demand specification with heterogeneous preference distributions given by equation (11) are presented in Table 5. Note that the results reported in Table 5 are based on the 40th and 35th income percentiles. Each column of the table contains the same set of control variables found in Table 2. The estimated income elasticities reported in Table 5 lie between 0.615 and 0.878 and are comparable to income elasticity estimates based on actual education capital spending in California of between 0.70 and 0.77 (Balsdon, Brunner and Rueben, 2003). The bottom panel of Table 5 reports the results of non-nested hypothesis tests based on the null hypothesis that the “correct” model is either the one based on the 50th and 45th income percentile difference or the 40th and 35th percentile difference. Note that in all four specifications, we reject the null hypothesis that the “correct” model is the one that includes the 50th to 45th income percentiles and fail to reject the null hypothesis that the “correct” model is the one that includes the 40th to 35th percentiles.²⁴ Furthermore, in the models where we utilize the 50th and 45th income percentiles, the estimated income elasticities are respectively 11%, 8%, 16% and 27% higher than the estimates reported in columns 1 through 4 of Table 5. That finding is consistent with the theoretical results of Goldstein and Pauly (1981) who demonstrate that in the presence of Tiebout sorting, income elasticities from standard median voter models are biased upwards by the heterogeneous preferences of voters.

The estimated coefficients on the variables we use to control for preference heterogeneity are generally negative and in many cases statistically significant. Specifically, in columns 1-4, the estimated coefficients on the variance of the fraction of low-demand voters and the Gini index are all negative and statistically significant. The coefficient on the racial heterogeneity index is also negative in columns 3 and 4 but never statistically significant. The negative sign on our preference heterogeneity controls is consistent with greater heterogeneity in the fraction of low-demand voters, income, and/or racial

²³ Results for counterfactuals using the 50th and 45th percentile incomes are quite similar.

²⁴ All results presented in the preceding paragraph arise in simple constant elasticity demand models that do not allow for heterogeneous distributions of preferences across districts.

heterogeneity leading to larger heterogeneity in preferences or greater variance in the unobservables that shape voting behavior as captured by β in equation 7. Specifically, our results suggest that the influence of the decisive voter's demand on the determination of voting rules is smaller in more heterogeneous districts. That finding is consistent with Romer, Rosenthal and Munley (1992) who demonstrate that greater preference heterogeneity within jurisdictions results in lower approval margins for local referenda. It is also consistent the empirical results of Gerber and Lewis (2004) who find that politicians tend to follow the median voter's preferences more closely in more homogenous jurisdictions.

{ Insert Table 5 Here }

To put our results into context, we can use the results reported in Table 5 to examine how a one standard deviation increase in the preference heterogeneity term in equation (11), $\exp(\hat{\phi} Z)$, scales the influence of the decisive voter's demand on the change in vote log-odds. Evaluated at the mean of $\exp(\hat{\phi} Z)$, a one standard deviation increase in the standard deviation of preferences leads to a 39.9% decrease in the contribution of the median voter's preferences to the log odds of voting yes in model 1, and a 29.8%, 34.0%, and 24.8% decrease in models 2, 3, and 4, respectively. In terms of the second referenda, a one standard deviation increase in heterogeneity in all school districts would have decreased predicted support for the referenda by 13.2 percentage points in model 1, and by 9.4, 7.8, and 14.4 in models 2, 3, and 4, respectively. Given that the actual percent yes in the second referenda was 54.4%, these predicted declines in support suggest that the referenda would have been defeated in all four models. It is also interesting to note that relative to models that use the 40th and 35th percentile incomes, models that use the 50th and 45th percentile incomes imply that a one standard deviation increase in preference heterogeneity leads to a larger predicted decline in the contribution of the median voter's preferences, particularly in models 3 and 4. Specifically, model 3 predicts a 43.3% (compared to 34.0%) decrease in the contribution of the median voter's preferences to the log odds of voting yes, and model 4 predicts a 47.8% (compared to 24.8%) decline.²⁵

²⁵ These larger predicted impacts are due to the fact that relative to the results reported in Table 5, models that use the 50th and 45th percentile incomes yield estimated coefficients on the preference heterogeneity variables that tend to be larger in magnitude. The largest difference is for the estimated coefficient on the variance of low demand voters in model 4 where the parameter estimate increases from -13.016 in column 4 of Table 5 to -23.522.

6. Heterogeneous Preferences and Pivotal Voter Income

Reduced Form Tests

Our empirical identification of the 40th income percentile as decisive under majority rule is consistent with the earlier findings of Fletcher and Kenny (2008) and Epple and Romano (1996a) that the income of the pivotal voter lies below the median when communities contain both high- and low-demand households. Specifically, our results are consistent with the hypothesis that low-income voters and relatively high-income voters with weak preferences for public services form a coalition to oppose the preferences of middle and high-income voters with strong preferences; a hypothesis that we shall refer to as the heterogeneous preferences hypothesis. To explicitly test that hypothesis, we used data from the special school district tabulations of the 2000 Census to calculate the fraction of all households in each district that are high income (incomes above \$75,000) and are likely to have low demand for public education spending. We focus on two groups of households who might be expected to have low demand: (1) household without children and (2) households with children in private school. We then used data on the fraction of high-income/low-demand households to split our sample into two equally sized subsamples. The first subsample contains districts in which the fraction of high-income/low-demand households is above the sample median and the second subsample contains districts in which the fraction of high-income/low-demand households is below the sample median. We then estimated separate regressions, similar to the “horse race” regressions reported in Table 3, for each of the two samples.²⁶

The results of that exercise are reported in Table 6, with the results based on subsamples of high-income households without children shown in the top panel and high-income households with children in private school shown in the bottom panel. Within each panel, the first set of rows contains estimates using income percentiles that are 5 percent above and below the 40th to 35th income percentiles selected in Table 3 (i.e. the 45th to 40th and the 35th to 30th income percentiles). The second set of rows contains estimates using income percentiles 5 percent above and below the 35th to 30th income percentiles, which in Table 3 yielded results that were statistically indistinguishable from the 40th to 35th percentile. The final set of rows directly compares the 40th to 35th and 35th to 30th income percentiles, which as noted above are virtually indistinguishable in the full sample. The first two rows in each set contain estimates for the subsample that contains school districts that have a fraction of high-income/low-demand households that is above the median for all school districts, while the last two rows contain estimates for the below median subsample. We hypothesize that the income percentile of the decisive voter should fall for the subsample of districts associated with the top row because these are the households expected to

²⁶ Fletcher and Kenny (2008) and many previous studies of demand for education have focused on the presence of older voters presumably because older voters are less likely to have children in local public schools. Our analysis for households without children attempts to capture this directly. Nonetheless, results dividing the sample based on share of households aged 55 or older yield similar results to those presented in Table 6.

vote with the poor against increased spending, and similarly the income percentile should rise for the bottom row.

Columns 1 and 2 of Table 6 present results based on “Horse Races” between the various income percentile differences. The results reported in those columns generally support the heterogeneous preferences hypothesis. Specifically, in each set of rows, the magnitude of the lower income percentile difference coefficient is larger for the above median subsample and the magnitude of the higher income percentile difference coefficient is larger for the below median subsample. The one exception is the “Horse Race” between the 40th to 35th and the 30th to 25th income percentiles in the top panel where the two estimates are very close in magnitude.

Similar to Table 3, columns 3 and 4 present the p-values associated with non-nested *P* tests that directly compare the performance of the two different income percentile differences listed in each set of rows. Specifically, column 3 reports p-values based on the null hypothesis that the “correct” model is the one that includes the higher income percentile difference listed in each set of rows, while column 4 reports p-values based on the null hypothesis that the “correct” model is the one that includes the lower income percentile difference listed in each set of rows.

The results reported in columns 3 and 4 once again generally support the heterogeneous preferences hypothesis. For example, in the top set of rows for the upper panel we find evidence in favor of the 35th to 30th percentile difference for the subsample that contains a higher fraction of high-income/low-demand households and evidence in favor of the 45th to 40th percentile difference for the subsample that contains a low fraction of high-income/low-demand households. More generally, if the heterogeneous preferences hypothesis is correct, significant p-values should fall on the diagonal of the 2x2 matrices on the right hand side of Table 6. This is generally what we find: 7 out of the 12 p-values on the diagonal of the 2x2 matrices are statistically significant and several others are near significance. Also note that none of the off diagonal p-values are close to being statistically significant. Results that are of particular interest are found in the bottom set of rows in both panels that compare the 40th to 35th percentile difference to the 35th to 30th difference. In table 3, these percentiles were indistinguishable in terms of changes in voting patterns. However, for the subsamples based on high income households with children in private school both tests reject the appropriate percentiles predicted by the heterogeneous preferences hypothesis. Similarly, for the subsample based on households with no children one test strongly rejects the appropriate percentile and the other test just misses rejecting the appropriate percentile at the 10 percent significance level.²⁷

²⁷ Subsample analyses similar to those presented in Table 6 were also estimated for constant elasticity demand models and finding are very similar, consistently supporting the heterogeneity in preferences hypothesis.

{Insert Table 6 Here}

Bimodal Distribution of Preferences with Constant Elasticity of Demand

In this section, we build on the work of Epple and Platt (1998) and Epple and Sieg (1999) by explicitly recognizing that there may be two types of voters in each community who are characterized by relatively low or high demand for education. In equilibrium under majority rule, the decisive voter must be the voter with median preferences overall, but within any group or type the decisive voter will likely not have the median income. In our model with two types, the typical low demand type will have a decisive voter with income above the median, and the typical high demand type will have a decisive voter with income below the median.

For each type of voter, individual preferences are distributed unimodally around the mean or mode preference for that type. Conditional only on income, the distribution of preferences is now bimodal because the overall distribution is a weighted composite of the distribution of preferences for low and high demand voters. Assuming that the income elasticity of demand is constant over income and type, the relative demand of low and high demand types can be described as

$$S_H(y) = b_H y^\theta = \eta S_L(y) = \eta b_L y^\theta, \quad (12)$$

where b_H and b_L are scalars and $\eta = b_H/b_L$. It follows that $\eta > 1$.

We continue to assume that conditional on income and demand type, individual voter preferences within a jurisdiction, μ_{ij} , follow independent type 1 extreme value distributions so that the share voting yes for a specific demand type is described by equation (7) as a function of the income of the decisive voter for a jurisdiction, demand type, and voting rule/referenda. Therefore, the change in vote share for a given demand type can be written as a function of the decisive voter's income for each referenda

$$\ln\left(\frac{pyes_j^2}{1-pyes_j^2}\right) - \ln\left(\frac{pyes_j^1}{1-pyes_j^1}\right) = d_i \exp[\phi Z_j] [y_{ji}^{1^\theta} - y_{ji}^{2^\theta}] + (\varepsilon_{ji1} - \varepsilon_{ji2}), \quad (13)$$

where i designates whether the demand is for high or low demand types, and y_{ji}^k is the decisive voter income for referenda k , jurisdiction j , and demand type i . Note that at the decisive voter income for each type, the demand of the two types must equal to ensure that all decisive voters support the same level of public service or output. That fact along with the fact that d_H equals ηd_L , implies that equation (13) is duplicated across types. This equal demand feature of our estimation equation is similar to Epple and

Platt (1998) and Epple and Sieg (1999) where the median voters form a locus of individuals over income and unobserved tastes for the public good.

The next two equations are defined so that the demand for each referenda in equation (13) is the optimal level of demand for the median preference voter or the 45th percentile preference voter in the case of the second referenda. Specifically, the decisive voter income percentile for each group determines the fraction of that group that votes against the referenda, and the average of the two groups' percentiles weighted by group size must equal the fraction required for passage or

$$(1 - P_k) = f_{jL}(y_{jL}^k)F_{jL} + f_{jH}(y_{jH}^k)(1 - F_{jL}), \quad (14)$$

where f_{ji} is the income distribution for type i in jurisdiction j , and F_{jL} is the share of low demand individuals in jurisdiction j . The share of voting age adults who are age 45 or older and do not have kids, as well as the income distribution for each group, is available from the school district files of the 2000 Decennial Census. The share low demand is then adjusted by the fraction of registered voters who are age 45 or older.

The decisive voter income for the high type in each referendum y_{jH}^k is then modeled as a function of the fraction of low demand voters and the logarithm of the ratio of the 60th to the 40th income percentile in the jurisdiction. A higher share of low demand voters requires that the high demand decisive voter income fall further in order to compensate for changes in the low demand decisive voter income. The income percentile ratio captures how dense the income distribution is between the 40th and 60th percentile. If the distribution is dense (low value on the variable), changes in the decisive voter income percentile lead to only small changes in income and thus small changes in the level of decisive voter demand implied by referendum passage. Consequently, with a dense distribution, a smaller percentile change in the high demand decisive voter's income is needed to adjust that decisive voter's demand to the demand of the low demand decisive voter. For the first referenda, the equation is structured to restrict the decisive voter percentile to be below the 50th percentile because more than half of the high demand voters must support a given demand if that demand is to satisfy the preferences of the median preference voter:

$$f_{jH}(y_{jH}^1) = 0.5 \cdot (e^{\lambda_1 X} / (1 + e^{\lambda_1 X})). \quad (15)$$

The equation for the second referenda takes a similar form except that it is restricted so that the high demand decisive percentile falls below the 45th percentile to assure 55 percent support for the public service demand level and to ensure that the percentile actually falls as predicted when the vote requirement is increased from 50 to 55 percent:

$$f_{jH}(y_{jH}^2) = \text{Min}(0.45, f_{jH}(y_{jH}^1)) \cdot (e^{\lambda_2 X} / (1 + e^{\lambda_2 X})). \quad (16)$$

In practice, the referendum 1 high demand decisive percentile always falls below the 45th percentile and is the binding requirement. Note that the decisive voter income for referenda 2 already depends upon the share of low demand voters in a jurisdiction through the presence of $f_{jH}(y_{jH}^1)$ in equation (16). We cannot identify an additional effect of this variable on the decisive voter percentile for referenda 2 because our model is differenced to eliminate jurisdiction fixed effects and thus we can only exploit information on the change in vote shares across referenda. However, the referenda 2 decisive percentile is allowed to depend upon the interaction of the fraction of low demand voters and the logarithm of the ratio of the 60th to the 40th income percentile. The drop in the high demand decisive voter percentile must match the decline in the low demand decisive voter percentile. While the change naturally depends upon the income distribution density, the magnitude of the effect on the high demand percentile will increase with the share low demand voters. Finally, the income of the low demand decisive voter for each referendum is found using equation (12) and requiring that the decisive voters for each type have the same demand in the voting equilibrium.

We estimate the parameters of equations 13 through 16 using a generalized method of moments (GMM) estimator. The GMM estimation model is estimated for the sample of jurisdictions and contains three moment equations, equation (13) and two versions of equation (14), one for each referenda. The instruments include the share of low demand voters, the share of low and high demand types in each of the 17 census income categories, the controls for modeling preference heterogeneity across jurisdictions, the linear controls in the log-odds voting equation, and the new controls for modeling the decisive voter income for the high demand type for each referendum.

Estimates are presented in Table 7. The income elasticity estimates range between 0.651 and 0.866 and are similar in magnitude to the elasticities reported in Table 5 for our simple constant elasticity of demand model. Thus, our results once again suggest that models that assume the median income voter is decisive tend to overestimate income elasticities when preferences are heterogeneous. The estimates of the preference heterogeneity parameters tend to be smaller in magnitude than those reported in Table 5, but continue to display significant effects for the variance associated with share of low demand voters and the Gini coefficient. The coefficient estimate on the variance associated with voters 45 years or older with no children cannot be directly compared to the estimate from Table 5 because some of the within community variation in preferences associated with this variable is captured by the two demand types and the bimodal preference distribution. However, we can compare the effect of changes in the Gini coefficient across the models presented in Tables 5 and 7. For models 3 and 4 respectively, the estimates

reported in Table 7 suggest that a one standard deviation increase in income heterogeneity as captured by the Gini coefficient leads to an 11% and a 6% decrease in the contribution of the median voter's preferences to the log odds of voting yes. Those predicted declines are smaller than the declines of 13% and 8% obtained from our simple constant elasticity model that utilizes the 40th and 35th percentile incomes and they are even smaller than the declines of 15% and 11% obtained from models that use the 50th and 45th percentile incomes. Therefore, the importance of income heterogeneity to voting outcomes is overstated in simple median voter models that rely on the median income of a jurisdiction to identify the decisive voter.

Turning to the parameter estimates for our decisive voter equations, we find that they generally match our predictions. Specifically, low income density increases the high demand decisive voter percentile for each referenda, and the share of low demand voters and the interaction of that share with the income density variable lowers the decisive voter percentile for the first and second referenda, respectively. In order to assure a positive η , we model the parameter as an exponential of the estimable parameter reported in the last row of Table 7. An estimate of zero yields a η of one implying equal demands across groups and is consistent with failing to reject the null hypothesis of one demand type. The estimates range from -0.541 to -0.596 and have very narrow 95 percent confidence intervals (0.03 or smaller). Thus, we strongly reject the null hypothesis of equal demands across groups. Specifically, our results imply the existence of a low demand type whose demand for the public good is between 0.55 and 0.58 percent of the high type's demand. For the model in column 4, these estimates are consistent with an average high demand decisive voter at the 35.3 percentile (\$38,128) for the first referenda and the 30.4 percentile (\$34,218) for the second. Similarly, the estimates are consistent with an average low demand decisive voter at the 70.8 percentile (\$69,343) for the first referenda and the 65.8 percentile (\$62,086) for the second. There is also considerable variation across communities in decisive voter percentiles. For example, for the first referendum the 95% confidence interval is between the percentiles of 28.6 and 42.0 for the high demand type and between 55.1 and 86.5 for the low demand type.

{Insert Table 7 Here}

7. Conclusion

This paper provides a direct test of the political economy “as if” proposition that underlies nearly all empirical studies that utilize the median voter model. We employ a unique dataset to examine whether the voter with the median income is decisive in local spending referenda. Previous tests of the median voter model have typically relied on aggregate cross sectional data to examine whether the voter with the median income is pivotal. These studies are likely biased because communities differ across a variety of unobservable dimensions that are likely correlated with the distribution of income in a community. In

contrast to previous studies, we make use of a unique natural experiment that allows us to estimate a first difference specification that controls for jurisdiction unobservables and avoids the fundamental problem of measuring the actual services demanded by voters. Consequently, we are able to avoid many of the problems that have hindered prior studies that have tested the median voter hypothesis.

Our empirical results suggest that voters understand the impact of changes in the identity of the decisive voter and rationally consider the impact of voting rules on local spending when voting on referenda that determine voting rules. The magnitudes of our findings appear to be quite reasonable and are consistent with previous literature. For example, our results suggest that the implied change in the decisive voter's income is consistent with between a 2.1 and 3.2 percentage point increase in the percent voting yes due to the change in the vote requirement from 50 to 55 percent, while the actual increase in percent voting yes was 4.3 percentage points. The constant elasticity of demand models provide estimated income elasticities of between 0.615 and 0.878, which are stable across specifications and consistent with the existing literature. Further, the estimated effect of median income on voting is not present in counterfactuals estimated at the census tract and state assembly district level.

However, our results strongly suggest that even under majority rule voting, the voter with the median income is not decisive. Rather, the results of both our linear model and our constant elasticity model are consistent with a decisive voter at the 40th percentile income for majority voting. That finding is consistent with the hypothesis that low-income voters and relatively high-income voters with weak preferences for public services form a coalition to oppose the preferences of middle and high-income voters with strong preferences. We directly test this hypothesis in two ways. First, we split the sample between districts that contain more or less high-income/low-demand households and all our findings support the hypothesis. We also estimate a model that allows for two distinct demand types defined by lack of children among older voters. The model estimates imply a strong rejection of a model that collapses to one demand type and finds that our low demand types (older voters with no children) have public service demands that are 45% lower than the high demand type. Further, the elasticity estimates of both the model with a decisive voter income below median and with two demand types are considerably below the income elasticity estimates arising from a simple median income voter model. That finding is consistent with Goldstein and Pauly's (1981) original conclusion that median income models will overstate income elasticities.

We also consistently find that school district income and demand heterogeneity are associated with reduced influence of the decisive voter's preferences on support for the referenda, a result consistent with Gerber and Lewis's (2004) analysis of politician's behavior. In terms of magnitude, we find that a one standard deviation increase in preference heterogeneity among all school districts would have reduced support for the second referenda (which passed by 4.4 percentage points) by between 7.1 and

12.6 percentage points. In addition, we find that the importance of income heterogeneity is overstated in models that ignore the existence of different groups of voters with distinct demand types. Earlier work by Romer, Rosenthal and Munley (1992) and Rothstein (1994) concludes that the decisive voter's preferences should have less influence on support for referenda in more heterogeneous jurisdictions, and our findings provide strong support for the implications of their theoretical models. Recent empirical work by Alesina, Baqir and Easterly (1999) and Alesina, Baqir and Hoxby (2004) finds that heterogeneous communities spend less on productive public goods and that jurisdiction consolidation is reduced when the surrounding region is heterogeneous, respectively. Our model along with the earlier work of Romer, Rosenthal and Munley (1992) and Rothstein (1994) identifies another important mechanism by which heterogeneity influences public choice concerning the provision of local public goods. Referenda models of this sort clearly imply that heterogeneity in preferences within a jurisdiction will reduce electoral support for both referenda that authorize spending on public services, as well as referenda intended to liberalize the rules under which spending is authorized.

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Table 1
Summary Statistics

Variable	Proposition 26		Proposition 39	
	Mean	Std. Dev.	Mean	Std. Dev.
<i>Difference Variables</i>				
Fraction Yes	0.477	0.085	0.505	0.084
Income	47,421	17,286	42,835	15,806
Turnout	0.279	0.098	0.431	0.122
Fraction Republican	0.373	0.114	0.374	0.114
Fraction Democrat	0.441	0.108	0.436	0.106
<i>Level Variables</i>				
		Mean	St. Dev.	
Fraction College Educated		0.224	0.150	
Fraction Homeowner		0.634	0.109	
Fraction H.H. White		0.640	0.217	
Fraction H.H. with Children		0.381	0.096	
Fraction Age 65 or Older		0.156	0.056	
Variance Voters 45 or Older No Children		0.234	0.015	
Gini Index of Income Inequality		0.423	0.039	
Herfindahl for Race/Ethnicity		0.489	0.147	

Notes : Table contains means and standard deviations for the variables used in the analysis. Summary statistics are based the sample of unified school districts in California. The income and tax price variables represent the 50th and 45th percentile values for Proposition 26 and 39, respectively.

Table 2
Coefficient Estimates: Linear Demand Model

	(1)	(2)	(3)	(4)
Decisive Voter Income (\$10,000's)	0.188** (0.052)	0.232** (0.066)	0.143** (0.059)	0.206** (0.092)
Turnout		0.220 (0.163)	0.182 (0.156)	0.102 (0.156)
Fraction Democrat		-0.804 (0.560)	-0.404 (0.428)	-0.191 (0.429)
Fraction Republican		0.052 (0.525)	-0.022 (0.464)	-0.231 (0.506)
Second Quantile of Size			0.019 (0.023)	0.018 (0.022)
Third Quantile of Size			0.085** (0.022)	0.081** (0.021)
Fourth Quantile of Size			0.105** (0.022)	0.094** (0.022)
Fraction College Educated				-0.114 (0.108)
Fraction Homeowner				0.046 (0.100)
Fraction H.H. White				-0.065 (0.057)
Fraction H.H. with Children				-0.273 (0.183)
Fraction Age 65 or Older				-0.554** (0.173)
R-Square	306	306	306	306
Observations	0.05	0.08	0.18	0.21

Notes: Columns 1-4 contain OLS parameter estimates for the change in log odds of share voting yes between the two referenda. The rows denoted by decisive voter income, turnout, fraction democrat, and fraction republican contain estimates on the change in those variables between the two referenda while the next eight rows contain estimates on the district size fixed effects and school district demographic attributes. Robust standard errors are shown in parentheses, and statistical significance at the 10% and 5% level are denoted by * and **, respectively.

Table 3
Coefficient Estimates: "Horse Race" Regressions

50 th - 45 th Percentile Income Difference Versus other Percentile Differences				
Percentile	Coefficient	St. Error	H ₀ : 50 th - 45 th	H ₀ : Other Percentile
50 th - 45 th	0.164	(0.114)	0.577	0.167
60 th - 55 th	0.061	(0.083)		
50 th - 45 th	0.179	(0.136)	0.817	0.176
55 th - 50 th	0.036	(0.115)		
50 th - 45 th	0.117	(0.165)	0.498	0.521
45 th - 40 th	0.118	(0.175)		
50 th - 45 th	0.076	(0.103)	0.023**	0.585
40 th - 35 th	0.244**	(0.108)		
50 th - 45 th	0.075	(0.104)	0.025**	0.591
35 th - 30 th	0.235**	(0.106)		
50 th - 45 th	0.128	(0.100)	0.147	0.31
30 th - 25 th	0.169	(0.112)		
40 th - 35 th Percentile Income Difference Versus other Percentile Differences				
Percentile	Coefficient	St. Error	H ₀ : 40 th - 35 th	H ₀ : Other Percentile
40 th - 35 th	0.264**	(0.107)	0.773	0.012**
55 th - 50 th	0.046	(0.085)		
40 th - 35 th	0.255**	(0.118)	0.750	0.028**
45 th - 40 th	0.046	(0.123)		
40 th - 35 th	0.173	(0.122)	0.207	0.183
35 th - 30 th	0.161	(0.122)		
40 th - 35 th	0.227**	(0.109)	0.428	0.048**
30 th - 25 th	0.109	(0.116)		
40 th - 35 th	0.212**	(0.102)	0.214	0.064*
25 th - 20 th	0.141	(0.102)		
40 th - 35 th	0.290**	(0.114)	0.982	0.007**
20 th - 15 th	-0.005	(0.105)		

Notes: Columns 1-2 contain OLS estimates, based on model 4 of Table 2, for various "horse races" between the difference in the 50th and 45th income percentiles (top panel) or the difference in the 40th and 35th income percentiles (bottom panel) and various other income percentile differences. In the top panel, columns 3-4 contain p-values from non-nested hypothesis tests (*P*-tests) based on the null hypothesis that the "correct" model is the model that includes the 50th and 45th income percentile difference or the other income percentile difference listed under the percentile column of Table 3. In the bottom panel, columns 3-4 contain the same information for various non-nested hypothesis tests that compare the 40th and 35th income percentile difference to various other percentile differences. Robust standard errors are shown in parentheses, and statistical significance at the 10% and 5% level are denoted by * and **, respectively.

Table 4
Coefficient Estimates from Counterfactuals

	(1) School Districts	(2) State Assembly Districts	(3) Census Tracts	(4) Census Tracts with District Fixed Effects
Model 1				
Income	0.226** (0.057)	0.013 (0.092)	-0.008 (0.019)	-0.020 (0.012)
Observations	306	80	6891	6891
Model 2				
Income	0.279** (0.071)	0.052 (0.089)	0.013 (0.017)	-0.009 (0.009)
Observations	306	80	6891	6891
Model 3				
Income	0.189** (0.065)	0.088 (0.094)	0.014 (0.017)	-0.008 (0.009)
Observations	306	80	6891	6891
Model 4				
Income	0.287** (0.093)	0.308 (0.370)	0.030 (0.019)	-0.017 (0.014)
Observations	306	80	6891	6891

Notes: Columns 1, 2, 3 and 4 present O.L.S. coefficient estimates for the difference between the 40th and 35th percentile income for the sample of school districts, state assembly districts, census tracts, and census tracts with controls for district fixed effects, respectively. The panels correspond to the models listed in Table 2 and all models contain the same control variables listed in Table 2. Robust standard errors are shown in parentheses, and statistical significance at the 10% and 5% level are denoted by * and **, respectively.

Table 5
Coefficient Estimates: Constant Elasticity of Demand with Preference Heterogeneity

	(1)	(2)	(3)	(4)
Income	0.753** (0.162)	0.853** (0.154)	0.878** (0.181)	0.615** (0.182)
Turnout		0.234 (0.151)	0.214 (0.147)	0.191 (0.153)
Fraction Democrat		-0.994** (0.366)	-0.611* (0.361)	-0.756** (0.376)
Fraction Republican		-0.120 (0.457)	-0.062 (0.466)	-0.163 (0.478)
Second Quantile of Size			-0.004 (0.023)	0.008 (0.022)
Third Quantile of Size			0.052** (0.021)	0.063** (0.022)
Fourth Quantile of Size			0.078** (0.022)	0.087 (0.023)
Fraction College Educated				-0.094 (0.100)
Fraction Homeowner				0.014 (0.111)
Fraction H.H. White				0.065 (0.050)
Fraction H.H. with Children				-0.200 (0.131)
Fraction Age 65 or Older				-0.404 (0.239)*
<i>Preference Heterogeneity Parameters</i>				
Variance Voters 45 or Older No Children	-19.799** (3.473)	-15.184** (4.007)	-16.401** (4.745)	-13.016** (4.099)
Gini Index of Income Inequality	-3.112** (1.394)	-2.775** (1.282)	-3.454** (1.535)	-2.268* (1.198)
Racial Index	0.567 (0.358)	0.235 (0.151)	-0.452 (0.435)	-0.257 (0.272)
<i>P</i> Test	p-value	p-value	p-value	p-value
H ₀ : 50 th - 45 th	0.006**	0.005**	0.015**	0.002**
H ₀ : 40 th - 35 th	0.830	0.956	0.926	0.632

Notes: Table presents the estimates from the constant elasticity of demand model with preference heterogeneity shown in equation (11). The estimates presented in the first row under income represents an elasticity while the other estimates are coefficients on the variables in a standard linear specification. The bottom panel of the table shows the results of non-nested *P*-tests based on the null hypothesis that the "correct" model is the model that includes the income percentile difference listed after H₀:. Robust standard errors are shown in parentheses, and statistical significance at the 10% and 5% level are denoted by * and **, respectively.

Table 6
Coefficient Estimates: Heterogeneous Preferences Hypothesis Regressions

High Income Households with No Children					
	Percentile	Coefficient	St. Error	H ₀ : Top Row Percentile	H ₀ : Bottom Row Percentile
Above Median	45 th - 40 th	-0.015	(0.091)	0.025**	0.942
	35 th - 30 th	0.242**	(0.113)		
Below Median	45 th - 40 th	0.613**	(0.209)	0.675	0.002**
	35 th - 30 th	0.125	(0.185)		
Above Median	40 th - 35 th	0.130	(0.108)	0.471	0.314
	30 th - 25 th	0.107	(0.151)		
Below Median	40 th - 35 th	0.711**	(0.234)	0.871	0.002**
	30 th - 25 th	-0.087	(0.192)		
Above Median	40 th - 35 th	0.050	(0.118)	0.108	0.744
	35 th - 30 th	0.199	(0.128)		
Below Median	40 th - 35 th	0.701**	(0.238)	0.88	0.003**
	35 th - 30 th	-0.053	(0.194)		
High Income Households with Children in Private School					
Percentile	Percentile	Coefficient	St. Error	H ₀ : Top Rown Percentile	H ₀ : Bottom Row Percentile
Above Median	45 th - 40 th	0.053	(0.115)	0.131	0.763
	35 th - 30 th	0.202	(0.138)		
Below Median	45 th - 40 th	0.325	(0.213)	0.305	0.128
	35 th - 30 th	0.228	(0.168)		
Above Median	40 th - 35 th	0.022	(0.126)	0.122	0.913
	30 th - 25 th	0.227	(0.154)		
Below Median	40 th - 35 th	0.526**	(0.229)	0.918	0.017**
	30 th - 25 th	-0.076	(0.169)		
Above Median	40 th - 35 th	-0.042	(0.137)	0.065*	0.848
	35 th - 30 th	0.268*	(0.149)		
Below Median	40 th - 35 th	0.458*	(0.261)	0.880	0.065*
	35 th - 30 th	0.051	(0.204)		

Notes: Columns 1-2 contain OLS estimates, based on model 4 of Table 2, for various "horse races" between different income percentile differences. Rows denoted Above Median correspond to the subsample of districts with high concentrations of high-income/low-demand households, while rows denoted Below Median correspond to the subsample of districts with low concentrations of high-income/low-demand households. Columns 3-4 contain p-values from non-nested hypothesis tests (*P*-tests). Specifically, column 3 contains the p-value based on the null hypothesis that the "correct" model is the model that includes the income percentile difference listed in the top row of each percentile comparison while column 4 contains the p-value based on the null hypothesis that the "correct" model is the model that includes the income percentile difference listed in the bottom row of each percentile comparison. Robust standard errors are shown in parentheses, and statistical significance at the 10% and 5% level are denoted by * and **, respectively.

Table 7
Coefficient Estimates: Bimodal Distribution of Preferences Model

	(1)	(2)	(3)	(4)
<u>Demand Equation</u>				
Income	0.818** (0.151)	0.876** (0.136)	0.861** (0.158)	0.695** (0.113)
<i>Preference Heterogeneity Parameters</i>				
Variance Voters 45 or Older No Children	-17.206** (3.580)	-11.742** (3.846)	-12.977** (4.660)	-8.653** (2.785)
Gini Index of Income Inequality	-2.473* (1.414)	-2.426** (1.231)	-2.684* (1.521)	-1.663* (0.842)
Racial Index	0.585 (0.448)	0.209 (0.392)	-0.465 (0.479)	-0.275 (0.216)
<u>High-Demand Decisive Income Equation First Referendum</u>				
Constant	0.455** (0.133)	0.735** (0.108)	0.487** (0.116)	0.799** (0.112)
Fraction Voters 45 or Older No Children	-3.166** (0.098)	-2.319** (0.086)	-3.166** (0.097)	-2.384** (0.089)
log(60 th Percentile Income / 40 th Percentile Income)	4.304** (0.323)	2.894** (0.207)	4.233** (0.292)	2.839** (0.209)
<u>High-Demand Decisive Income Equation Second Referendum</u>				
Constant	1.189** (0.097)	1.451** (0.095)	1.203** (0.097)	1.478** (0.092)
log(60 th Percentile Income / 40 th Percentile Income)	1.926** (0.275)	0.986** (0.269)	1.894** (0.275)	0.937** (0.264)
Interaction of Fraction Voters 45 or Older No Children and log(60 th Percentile Income / 40 th Percentile Income)	-0.848** (0.232)	-0.028 (0.208)	-0.847** (0.229)	-0.039 (0.210)
<u>Low-Demand Decisive Income Equation</u>				
kappa=1/exp(η)	-0.596** (0.014)	-0.553** (0.012)	-0.595** (0.014)	-0.541** (0.012)

Notes: Table presents GMM estimates from the model shown in equations (12) through (16). The estimates presented in the first row under income represents an elasticity while the other estimates are coefficients on the variables in a standard linear specification. Robust standard errors are shown in parentheses, and statistical significance at the 10% and 5% level are denoted by * and **, respectively.