The Effects of Community Heterogeneity on Revealed Preferences for Environmental Goods

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Abstract

This paper examines the effects of community heterogeneity on the public provision of environmental goods. Analysis of 2000 referenda data of California census-tracts shows that more diversity is correlated with lower support for environmental initiatives. Although whites and Hispanics are more likely to support environmental initiatives only as the shares of their groups increase, there are negative effects of homogeneity among African-Americans and Asians. Regarding education, the more homogenous the community is the greater the support for the environmental initiatives. However, this result is driven by the high-school dropout group.

1. Introduction

Economists have long examined election results in order to understand preferences for public goods (see Deacon and Shapiro, 1975; Fischel, 1979; Matsusaka, 1992; or Dubin et al., 1992, among the seminal papers in this literature). The analysis of voting on referenda is particularly helpful to understand the determinants of the demand for environmental quality (Khan and Matsusaka, 1997; Khan, 2002). The natural environment is a public good and willingness to pay must be inferred either through direct methods (stated preferences) or indirect methods such as hedonic price and averting expenditures models. Voting on referenda is another way to study revealed preferences. Its main advantage is that it mimics decision-making in private markets. The election process allows voters to learn about the public good and its income and price effects. Voters will support the supply of a public environmental good if their willingness to pay is equal to or greater than the cost.

The literature on revealed preferences for environmental goods has adopted the view that a voter considers the effects of her decision on her welfare only. This standard approach controls for income and identifies variables that proxy the "prices" the voter faces, for example, occupation and educational attainment. Researchers also include race or ethnicity if the public good might create differences in well-being among different groups, for example if minorities are more exposed to environmental pollution. Empirically, in order to test the hypothesis that the race or ethnicity of the median voter matters, for example, researchers include the share of the population in a community that is of a given race or ethnicity. In this view, socio-demographic characteristics simply change the profile of the median voter across communities.

However, willingness to sacrifice income and the value attached to environmental amenities might depend on who benefits and who pays the cost of the public good other than the voter herself. In other words, socio-demographic factors might not only change the profile of the median voter but also influence the incentives of a voter to provide the environmental goods.[1] Indeed, recent work shows that increased ethnic and social-class diversity affects public spending (Alesina, Baqir, and Easterly, 1999) and individuals' efforts to engage in voluntary organizations (Alesina and La Ferrara, 2000) and undertake actions with public benefits (Vigdor, 2004; Okten and Osili, 2004). Typically, papers in this literature test for the effects of diversity

[1] Cutler, Elmendorf, and Zeckhauser (1993) distinguish between a self-interested model of public spending and a "community preference" model. The authors infer evidence for the community-preference model via the effects of characteristics of surrounding areas on a jurisdiction's spending.

via an index of community heterogeneity that is calculated as one minus the sum of group shares squared and is interpreted as the probability that two individuals selected at random belongs to different groups (Easterly and Levine, 1997).

This paper connects these two areas of the literature: the demand for environmental goods through revealed preferences in referenda and the effects of diversity on the provision of public goods. The paper investigates whether the social context in which the median voter makes her decision influences support for two environmental initiatives in California. Analysis of 2000 census-tract data shows that community racial and social-class diversity (as measured by educational attainment) influences the demand for environmental goods. Although more diversity is consistently correlated with lower support for environmental initiatives, there are differences in how racial homogeneity affects election results. While whites and Hispanics are more likely to support environmental initiatives the more homogenously white and Hispanic the community is, there are negative effects of homogeneity among African-Americans and Asians. These qualitative differences across racial groups and the fact that the models include voter turnout as a control for civic behavior suggest that the findings cannot be uniquely attributed to generalized attitudes toward civic behavior and that there might exist differences in preferences across racial groups for the public goods that the ballots proposals provide. Greater homogeneity can lead to greater coordination and lower demand for environmental regulation if one group understands the provision of the public good to be a "bad." The results also show that the more homogenous the community is in terms of educational attainment, the greater the support for the environmental initiatives. However, this finding is driven by the high-school dropout group. Finally, the paper presents evidence that suggests endogenous sorting is likely to bias diversity effects down so that the parameter estimates of the models could be interpreted as a lower bound of the actual effects. Although the paper discusses results of linear probability models, the results are robust to the estimation method.

Exploring the role of community heterogeneity is important to understand how institutional and social factors that affect the socio-demographic composition of a community might facilitate or hinder the public provision of environmental goods. The results of this research suggest that support for pro-environment initiatives varies across socio-demographic groups. Policymakers might consider to what extent the successful implementation of policies depends on recognizing and addressing the fact that environmental concerns are not shared equally by different socio-economic groups.

The paper is organized as follows. Section 2 summarizes relevant papers in each literature. Section 3 presents the environmental initiatives that are the object of study and two non-environmental ballot proposals that are examined for comparison purposes. Section 4 discusses the variables and specifications. Two sets of models are estimated. A first set of models follow closely the specification presented by Khan (2002). These base models are expanded to control for racial and social-class homogeneity. Section 5 presents results and explores the potential bias due to endogenous sorting. Section 6 concludes.

2. Related Literature

This paper connects two areas of the economics literature. One examines the demand for environmental goods through revealed preferences in referenda. The second literature explores the effects of community heterogeneity on the provision of public goods.

In a seminal paper, Deacon and Shapiro (1975) examine voting results for two initiatives in California in the early '70s, one to conserve the California coastline and the other to provide public financing for rapid transit. The authors use data for 344 California cities and find that education, political preference, and income were consistent determinants of support for the initiatives. Fischel (1979) is an early study of individual voter responses in eight New Hampshire towns. Fischel finds that income, occupation, and education are significant predictors of support for allowing a wood pulp processing mill to locate in a community.

Matsusaka and Kahn (1997) analyze the cross-county determinants of support for environmental ballot propositions in California between 1970 and 1994 focusing on income and price effects of the propositions. To proxy for the price of the initiatives, Matsusaka and Kahn use employment in different industries and occupations, educational attainment, and urban population. Matsusaka and Kahn argue that income and the proxies for price are sufficient to explain the demand for environmental protection and that non-economic factors such as ideology and political preferences add little explanatory power. Kahn (2002) expands this research by analyzing six additional California ballot proposals between 1994 and 1998 using data at the census-tract level. His explanatory variables include the percentage of the tract's population over 65, percentage black, percentage Hispanic, percentage of college graduates, median household income, density, and share of resident who work in different industries. Kahn finds that the more educated, the young, minorities, and those not working in polluting industries are more likely to demand environmental quality.

These studies provide insights about the profile of the median voter and average preferences for public environmental goods. However, heterogeneity can reduce social interaction and involvement and hinder the coordination of preferences. Heterogeneity can also influence a voter's decision if a ballot proposal has local benefits that would accrue to people from different groups. To investigate this hypothesis this paper draws from the literature that explores community heterogeneity effects.

Researchers have identified at least two overlapping channels through which diversity can matter. The socio-economic composition of a community determines the costs of engaging in collective action and deciding on the provision of public goods. If individuals from different groups hold different preferences, this diversity of preferences can polarize a community and result in the underprovision of public goods. A similar effect can occur if individuals suffer a loss of utility when interacting with people of different socio-economic backgrounds since the provision of public goods through collective action requires generalized local involvement via successful leadership and trust (Okten and Osili, 2004). For example, Alesina and La Ferrara (2000) study the links between community composition and social interaction (measured by affiliation with voluntary organizations) using data from the General Social Survey for the years 1974-1994. The authors find that diversity has negative effects. Membership in groups that require more interaction (such as a church group) is most affected by community fragmentation.

A second channel for diversity effects is differential altruism. People can value differently how their contributions to a public good affect their own and other groups in the community. Individuals might undervalue public goods that would accrue to different groups and value those goods that would be enjoyed by people from their own group. Vigdor (2004) examines how heterogeneity across racial, social class, and age groups affects the propensity to return the Census questionnaire, an action that has public benefits because the distribution of federal grants depends on Census population counts. Vigdor finds that counties that are more homogeneous have higher Census response rates. In addition, Vigdor finds that that black and Hispanic households and households with college degree are particularly responsive to the share of their own groups in each county. Following Vigdor (2004), this paper presents results from one specification that includes an index of homogeneity and results from a second specification that allows for the effects of homogeneity to be group-specific by decomposing the homogeneity index into each of its components (Vigdor, 2004, 2002).

3. California Initiatives

Research on revealed environmental preferences through voting on referenda has mostly focused on ballot proposals in California. California provides a unique opportunity to understand voting patterns because of its environmentalism and frequency of initiatives. Although its size and diversity imply that results might not be easily generalized to the U.S. at large, substantial variability within the state facilitates the task of estimating the factors driving the demand for environmental regulation.

This paper examines two environmental initiatives in the March 2000 primary elections: ballot proposals 12 and 13. Proposition 12 provides for a bond issue of $2,100 million dollars to buy, develop, and protect natural areas for recreational purposes. The acquisition of land and the development and protection of natural areas could affect the construction industry and forestry and farming workers. Some benefits would accrue to city dwellers since the part of the bond funds is used for recreational areas in urban areas. The California Chamber of Commerce supported the proposition citing its potential befits for tourism. The Chairman of the Black Chamber of Commerce of Los Angeles County was one of the signatories of the rebuttal to the argument in favor of the proposition. Proposition 13 provides for a bond issue of $1,970 million dollars to drinking water facilities, flood and watershed protection, water pollution treatment and control, water conservation, and water supply reliability. The acquisition and restoration of land for flood protection and watershed protection could affect the construction industry. The proposition would provide loans and grants to local agencies for water pollution control and restoration. This might explain why a diverse group of organizations supported the measure, including the Agricultural Council of California, the California Manufacturers Association, and the California State Council of Laborers. Ballot proposal 12 passed with 63.2 percent of the total votes in favor and ballot proposal 13 with 64.8 percent of the votes.

Following Kahn (2002), this paper explores two 2000 non-environmental initiatives for comparison purposes: ballot proposals 22 and 25. Ballot proposal 22 adds a provision to the California Family Code to the extent that the state only recognizes marriage between a man and a woman. This initiative has no fiscal effects on state and local governments. Ballot proposition 25 proposes campaign reforms: it expands the requirements for contribution disclosure, sets contribution limits, bans corporate contributions, and limits the period for fund-raising. This initiative was estimated to cost about $55 million annually as the state government would provide larger public funds for campaigning. Ballot proposal 22 passed with 61.4 percent of the votes and initiative 25 failed with only 34.7 percent of the votes in favor.

This paper does not attempt to test specific hypotheses about initiatives 22 and 25. Rather, the models that are used for the environmental proposals are estimated for these social and political issues. The goal of this exercise is to explore whether the independent variables have different qualitative and quantitative effects across initiatives in expected ways. Naturally, the distributional impacts of ballot proposition 22 depend on sexual orientation. Although the models could include the proportion of households with same-sex unmarried partners, same-sex couples are typically concentrated in urban areas and more racially diverse communities (Gates and Ost, 2004). Thus, some of the independent variables will account for a community's propensity to vote against or in favor of proposal 22. Regarding proposition 25, although it is also a controversial issue, it is difficult to propose informed hypotheses about the influence of socio-economic factors on support for campaign reforms. A priori, limits and bans on contributions might influence individuals in different industries and social classes differently to the extent that these variables account for the presence of special interests. It is expected that the models will explain well variability in initiatives 12 and 13, probably less well proposition 22, and explain poorly voting for proposition 25.

The proportion of votes in favor of proposal 12 is highly and positively correlated with the proportion of votes in favor of proposal 13 (the correlation coefficient is .97). The correlation coefficients between votes in favor of proposal 12 and proposal 13 and votes for limiting same-sex marriage are negative and fairly high at -.66 and -.61, respectively. Ballot proposal 25 is an interesting ballot for comparison purposes as voting patterns for this initiative and proposals 12, 13, and 22 are weakly correlated. Data on election results for these ballots proposals come from Institute for Governmental Studies at the University of California Berkeley. Information about the initiatives comes from the California Ballot Measures Database (University of California, Hastings College of the Law).

4. The Empirical Model

This paper presents two sets of models. First, in order to report a set of comparable results, the paper discusses results of models that follow the specification presented by Khan (2002). These base models are expanded to include homogeneity indices and group shares squared for race and educational attainment. The tables present R-squared values, log-likelihoods, and the Bayesian information criterion (BIC) to compare the expanded models to the base models.

Following Kahn (2002), the unit of analysis is census tracts.[2] In all cases, the dependent variable is the share of tract voters who voted in favor of a given ballot proposition (Table 1 presents summary statistics). In addition to median household income, the base models include the proportions of workers in manufacturing (PCTMANUF), farming or mining (PCTAGRI), and in finance, insurance, and real state (PCTFIRE). These variables control for the potential economic loss for workers in these industries. Population density (DENSITY) controls for differential benefits between rural and urban areas. Population density squared controls for non-linear benefits of public goods due to congestion. The variable PCTPOP65 measures the percentage of the population at least 65 years old. Although younger people are usually believed to be more pro-environment, the elderly are likely to suffer more from environmental than younger individuals. Kahn (2002) also points out that support for environmental regulation might vary over the life-cycle if middle-aged individuals are less likely to vote in favor of initiatives that impose a tax burden. The shares of Hispanics, African-Americans, Asians, and Native Americans in the population of each tract (PCTHISPANIC, PCTBLACK, PCTASIAN, PCTNATIVE) might matter if minorities are more exposed to environmental problems than whites are. The proportion of individuals over 25 years of age who have at least a college degree (PCTCOLLEGE) can control for effects due to variability in skills within specific occupations and for the potential influence of greater knowledge of the ballot propositions. The models also include the proportion of high-school graduates, PCTHS, so that the default educational group is high-school dropouts. Data for all these variables come from the Census 2000 Long Form files.

There are a few departures from the model estimated by Khan (2002). First, the models in this paper include the percent of votes for Republican candidates in the 2000 primary election,

[2] According to the Census, tracts are defined by permanent visible boundaries (streets, roads, rivers) and contain between 2,500 and 8,000 residents.

PCTREP. This variable improves the goodness of fit of the models and avoids biases in the coefficients on shares of minority population. Hispanics and blacks in California are less likely to vote Republican than whites are (the correlation coefficients are around -.50). If the models did not include political preference then the average effect of being Hispanic could be confounded with the average effect of voting Democrat (these findings are explained in the following sections). Second, the models include voter turnout (data for this variable and vote for Republican candidates come from Institute for Governmental Studies at the University of California Berkeley). To the extent that homogeneity is correlated with generalized trust and attitudes toward civic participation, it is possible that the estimated effects of homogeneity would in part capture the importance of civic behavior on the provision of public goods.[3] To control for civic behavior at the tract level, the variable TURNOUT is calculated as total votes divided by voting-age population.[4] Third, these models include income squared. Adding this variable improves the goodness of fit and the turning points fall within the sample distribution (although at the extreme of the distribution).

The expanded models include all the variables discussed above and also control for community heterogeneity. Following Vigdor (2004), this paper considers five racial groups (non-Hispanic white, non-Hispanic black, non-Hispanic Asian, non-Hispanic Native American, and Hispanic) and three education groups (high-school dropouts, high-school graduates, and college graduates). Two specifications are estimated for each ballot proposal and dimension of diversity. One specification includes an index of homogeneity calculated as the sum of group shares squared (and is equal to 1 minus the fragmentation index).[5] The second specification decomposes the homogeneity index into each of its components by estimating the coefficients on group shares squared. Using the index, the researcher estimates the average effect of homogeneity. By decomposing the index one allows for the effects of homogeneity to be group-specific (Vigdor, 2002). In other words, one estimates two parameters for each group (one intercept and one slope

[3] The coefficient on voter turnout should not be interpreted causally as it is possible that it is the decision to vote on a given ballot proposal that influences turnout.

[4] This calculation underestimates actual voter turnout because voting-age population includes individuals who are illegible to vote. However, the results are unchanged if the specifications include the percent of non-naturalized foreign-born population in the tract. In addition, county fixed effects will capture some of the possible effects of differences in foreign-born population across counties.

[5] The homogeneity index equals one in a perfectly homogenous community, for example, all residents are Hispanic, and approaches zero if the community is composed of many small groups. The fragmentation index, one minus the homogeneity index, is usually interpreted as the probability that two randomly selected individuals belong to different groups (Easterly and Levine, 1997).

for each group). A priori, it is interesting to allow for group-specific homogeneity effects some groups believe the ballot proposal would provide a public "bad" rather than a public good. For example, while the California Chamber of Commerce supported proposition 12 because of its potential benefits for tourism, the Black Chamber of Commerce of Los Angeles County opposed the same proposition.

All the models include county fixed effects (the tables do not report these coefficients). Linear probability models (weighted by the population of each tract), robust regression, median regression, and grouped logit regression were used to estimate the models. The results are very similar qualitatively.

5. Results

Table 2 reports the results of the models that follow a standard specification (similar to Kahn, 2002). Table 3 and Table 4 show the results when the models include homogeneity indices and group shares squared, respectively. For the linear probability models, the log-likelihood reaches a maximum when the density is greater than 1, thus the log-likelihood in these models is positive and the BIC is negative. The tables present parameter estimates of linear probability models. For initiatives 12, 13, and 22, less than 1 percent of the predicted probabilities fall outside the range of 0 to 100 percent while all predicted probabilities fall within the 0 to 100 range for initiative 25. Because the results for weighted least-square grouped logit estimations are qualitatively very similar and the estimates of the linear probability model are easier to interpret, the latter are the results discussed here. To account for heterocedasticity, the models are weighted by the population of each tract and robust standard errors are computed.

In terms of goodness-of-fit, the models explain around 90 percent of the variability in support for the two environmental initiatives. As one would expect, the models for limits-on-marriage and campaign reform explain a lower proportion of the variability in the support for these initiatives (this result is particularly evident for ballot proposal 25 with an R-squared of approximately. 50). Although including the homogeneity indices or group shares squared add little to the R-squared of the models, likelihood-ratio tests indicate that there is overwhelming statistical evidence that the specifications that include either group shares squared or indexes of homogeneity fit the data better than the comparable base models. For example, comparing the log-likelihood of column 1 in Table 3 to the log-likelihood of column 1 in Table 2, there is overwhelming statistical evidence that a model that includes the racial homogeneity index fits the data better than the base model (the chi-square statistic with 1 degree of freedom is 54.22 with a p-value below .1 percent). Thus, controlling for racial homogeneity improves the model. This conclusion also holds when accounting for educational attainment homogeneity. For example, comparing the model in column 2 of Table 2 to the model in column 4 of Table 4, a likelihood ratio test produces a chi-square statistic with 3 degrees of freedom of 110.2 with a p-value below .1 percent. The values of the BIC also indicate that even after penalizing the model for the additional explanatory variables, models that include the indices or group shares squared are statistically better than the base models (because the log-likelihood is positive for these models, the largest negative number for the BIC indicates the best model).

In addition, in some models the point estimates of group shares change signs when the models include group shares squared indicating that some coefficients might be biased when the model does not account for group-specific homogeneity effects. More importantly, decomposing the index into group shares squared provides interesting insights about which particular groups are driving the diversity effects. These results are discussed in detail in sections 5.2 and 5.3. Overall, models that account for the degree of homogeneity in a community produce results that are statistically better and substantively richer than the base models.

5.1 The Base Models

The results in Table 2 that are consistent across all models are explained first. Sections 5.2 and 5.3 discuss the coefficients for race and educational attainment. The coefficients of income and income squared reveal a convex relationship between median household income and support for the environmental initiatives. These estimates are statistically significant at the 1 percent level or better. For example, the coefficients in column 1 and column 2 indicate that the turning point occurs when median household income is approximately $100,000. This result is consistent with Kahn (2002) who argues that the public provision of environmental goods, rather than the goods themselves, might be considered an inferior good as income increases. There is also a convex relationship between income and support for proposition 25 but the relationship between income and support for proposition 22 is concave with a turning point at a higher level of income (approximately $200,000). The 99-percentile of the median household income distribution is approximately $135,000 and the maximum value observed in the sample is $200,001.[6]

Also consistent with Kahn (2002) and Khan and Matsusaka (1997), opposition to the environmental initiatives is stronger in tracts with more employment in farming and mining. This result indicates that opposition comes from those sectors that are more likely to be negatively influenced by the propositions. Employment in manufacturing has a positive and significant effect for initiative 13 (the California Manufacturers Association supported this water bond) while employment in industries where workers are likely to be more educated is also positively correlated with support for these initiatives. The coefficients on PCTMANUF, PCTAGRI, and PCTFIRE are all positive and statistically significant in the model for initiative 22. This finding suggests that higher employment in the omitted industries (such as the service industry) is negatively correlated with support for the ban on same-sex marriages. In the model for initiative 25, the parameter estimates for PCTMANUF and PCTAGRI are negative, possibly measuring the importance of special interests in these sectors.

The proportion of the population 65 years of age and older is positively correlated with votes for the environmental proposals, however, the effect is statistically significant for initiative 13 (water quality) but not for initiative 12 (parks and water). PCTPOP65 is also positively related to support for the initiative that imposes limits on marriage but is negatively related to the support for campaign reforms.

The coefficient on density is positive and statistically significant in the models for initiatives 12 and 13. This result can be attributed to the fact that the benefits created by these propositions are likely to accrue to city dwellers. There is evidence of a non-linear effect in the model for 12. Density has a weak statistically positive effect on support for initiative 22 but has a negative effect on support for ballot proposition 25. This finding suggests that individuals in urban populations were less likely to support limits and bans on campaign contributions that individuals in rural areas, everything else equal.

The effect of political preference is very strong substantively and statistically.[7] Communities that vote Republican are less likely to support the environmental initiatives and

[6] All models were also estimated via robust and median regression to control for potential outliers and influential observations. The results in terms of statistical significance and substantive findings were very similar to the results of the linear probability model.

[7] The R-squared of the base models decreases by approximately .10 for the environmental initiatives and ballot proposal 22. For initiative 25, the R-squared drops by approximately .02 only.

limits on campaign contributions than communities that vote Democrat do. On the other hand, communities that vote Republican are more likely to support the ban on same-sex marriage than communities that vote Democrat do. Excluding political preference biases the estimates on some race group variables. When political preference is not in the model, the results indicate that there is opposition to the limits-on-marriage proposition among Hispanics but the coefficient on Hispanic is positive and significant when political preference is in the model. The importance of political preference for the propositions in this analysis stands in clear contrast to the minor impact that such factors have in the analysis by Kahn and Matsusaka (1997). Besides the fact that the unit of analysis is different, the 2000 elections were a period of political polarization. Voter turnout is negatively correlated with votes in favor of ballot propositions 12, 13, and 25 but positively correlated with support for limits on marriage. This pattern of effects aligns with the estimated coefficients on the percent of votes for a Republican candidate.

Quantitatively speaking, for ballots proposals 12, 13, and 22 the largest effects are due to percent Hispanic, educational attainment, political preference, and median household income. For the proposal on campaign reform, the largest effect is due to median household income. Finally, the independent variables have different qualitative and quantitative effects across ballots proposals in expected ways. For example, higher educational attainment correlates with more votes in favor of the environmental initiatives and with fewer votes for the limit-on-marriage proposal and those who vote for a Republican candidate are more likely to vote for the limit-on-marriage initiative and against the environmental initiatives.

5.2 Racial Homogeneity

Table 3 reports the parameter estimates of the linear probability models that include either the homogeneity index or group shares squared. Consistent with the literature, racial homogeneity is positively correlated with support for the provision of public goods. The coefficients on the homogeneity index in the models for the two environmental initiatives are positive and significant at the 1 percent level or better. The coefficient estimates are such that one standard deviation increase in the index increases by .044 standard deviations the proportion of votes for initiative 12, everything else equal.[8] For comparison purposes, one standard

[8] The magnitudes of the effects due to changes in the indices are somewhat misleading as it is impossible to increase the index without changing the values of the group shares.

deviation increases in PCTFIRE and PCTCOLLEGE increase support for initiative 12 by .016 and .37 standard deviations respectively, while one standard deviation increases in PCTAGRI and PCTREP reduce support for initiative 12 by .054 and .841 standard deviations respectively. Racial homogeneity is also positively correlated with support for ballot proposal 22. One standard deviation increase in the index increases the proportion of votes for the limits-on-marriage initiative 22 by .093 standard deviations. This finding is probably due to the fact that same-sex couples tend to live in more racially diverse communities (Gates and Ost, 2004). It is less clear why more ethnically homogenous communities are less likely to support ballot proposal 25.

Decomposing the homogeneity index into group shares squared offers interesting insights. First, there is evidence that the coefficients of percent black and percent Hispanic in the model that do not allow for group-specific effects are biased. In the first specification, these coefficients are negative (although statistically not different from zero). In the model that includes group shares squared, the coefficients are positive and significant. Thus, consistent with Khan (2002), communities where minorities are a greater fraction of the population are more likely to support the public provision of environmental goods than whites do. Second, while whites and Hispanics are more likely to support these initiatives as their share in the population increases, there are negative effects of homogeneity for African-Americans and Asians. The effects for whites and Hispanics are what we expect under the hypothesis that ethnic diversity reduces social interaction or implies differential altruism. The results for African-Americans and Asians might appear counterintuitive. However, there was opposition to these initiatives among the African-American business community. In particular, the Black Chamber of Commerce of Los Angeles County was one of the signatories of the rebuttal to the argument in favor of ballot proposition 12. Greater coordination and within-group homogeneity can lower the demand for environmental regulation if one group understands the provision of the public good to be a "bad".[9] These results show the benefits of decomposing the homogeneity index and suggest that researchers and policy-makers can gain interesting insights from this modeling approach.

[9] There are also statistically significant negative homogeneity effects on the Asian and African-American group shares squared variables in a sample that excludes all census tracts in Los Angeles County.

5.3 Educational Attainment Homogeneity

Table 4 reports the parameter estimates of the models that control for diversity in educational attainment. Regarding the coefficients on group shares, the results are consistent with previous literature: communities where a larger share of individuals holds college degrees support the public provision of public goods more than communities where educational attainment is lower, everything else equal. Higher educational attainment is negatively correlated with support for the ban on same sex-marriage and limits on campaign contributions.

The more homogenous the community is in terms of educational attainment the greater the support for the environmental initiatives. As the index increases by one standard deviation, the proportion of votes for the environmental initiatives increases by approximately .03 standard deviations, everything else equal. However, the models with group shares squared indicate that these results are due to positive homogeneity effects in the high-school dropout group. The coefficients on group shares squared for college degree and high-school degree are statistically insignificant but the coefficients on the high-school dropout group share squared is positive and significant at the 1 percent level or better (Table 4, columns 2 and 4). This finding indicates that, after controlling for household income, occupation, and race, members of the most economically disadvantaged group are more likely to support the public provision of environmental goods as their share of the tract population increases.

Decomposing the index into group shares squared offers a different pattern of effects for initiatives 22 and 25. Although support for the ban of same-sex marriage is lower in communities where more individuals hold college degrees, as the share of this group increases support for the ban also increases. The opposite effects are found for the high-school group. Regarding initiative 25, support for limits on campaign contributions is lower in communities where more individuals hold college degrees and as the share of this group increases opposition to this initiative becomes stronger.

5.4 Check for Endogenous Sorting Bias

A common limitation of studies that explore heterogeneity effects is the possibility that the demographic composition of the unit of analysis is not random as individuals endogenously sort into communities.[10] Endogenous sorting is likely to bias estimates downward since people who move to more heterogeneous communities because of their preference for diversity are also less likely to oppose the public good as diversity increases.

Following Vigdor (2004), who argues that it is unlikely that individuals choose the county in which to live based on the socio-demographic composition of the county, the models were estimated using county-level data. However, given the complexity of the models and the fact that there are only 58 counties in California, the coefficients were estimated rather imprecisely and 95 percent confidence intervals for the homogeneity indices were notably wide.

A second approach is to split the sample into three sub-samples: tracts below the 25 percentile of the homogeneity index, tracts in the interquartile range, and tracts above the 75 percentile of the homogeneity index. If there is a positive correlation between a taste for diversity and the public provision of environmental goods, the coefficients of group homogeneity should be smaller (in absolute value) in the less homogeneous samples.[11] Tables 5 through 8 present the results for the two environmental initiatives.

The results in Table 5 show that the coefficients on African-American and Asian group shares squared are larger (in absolute value) in the most homogenous sample, as it would be expected if there were endogenous-sorting bias. On the other hand, the coefficients on the Hispanic group share squared are larger in the less homogenous samples and the parameter estimates of the index offer no conclusive evidence (the estimate of racial homogeneity in the interquartile range is larger than in the 25-percentile but lower than in the 75-percentile). Table 6 shows very similar results for initiative 13. The results for educational homogeneity are clearer. Table 7 shows that the homogeneity index is the lowest in the sample of least homogenous tracts and the parameter estimate of the high-school dropout group share squared is the largest in the most homogenous communities. The results in Table 8 are similar. Overall, these findings, particularly for homogeneity in educational attainment, show evidence that the parameter estimates of the indices and group shares squared found for the entire sample of tracts are probably a lower bound of the actual effects.

[10] As Alesina, Baqir, and Easterly (1999) argue, some degree of community heterogeneity is likely to persist because of legal constraints, economies of scale in the provision of public goods, limited mobility, and the multidimensional features of most public goods.

[11] For the index of racial homogeneity, the 25-percentile is .49 and the 75-percentile is .75. For the index of education homogeneity, the 25-percentile is .32 and the 75-percentile is .38.

6. Conclusions

This paper examines the effects of diversity along racial and educational attainment lines on the public provision of environmental goods. Analysis of 2000 census-tract data of election results of two environmental initiatives in California shows that more diversity is consistently correlated with lower support for environmental initiatives. Interestingly, there are differences across groups in how homogeneity affects election results. While whites and Hispanics are more likely to support environmental initiatives in less heterogeneous communities, there are negative effects of homogeneity among African-Americans and Asians. Regarding education, the more homogenous the community is the greater the support for the environmental initiatives. However, the models with group shares squared indicate that these results are due to strong affinity effects in the high-school dropout group.

In summary, this paper finds that the degree of racial and educational attainment homogeneity in a community influences the demand of environmental goods through voting on referenda and that decomposing homogeneity indices produces a richer picture of the links between socio-demographic characteristics and the demand for environmental goods. These results can be useful for researchers, advocates, and policymakers studying and implementing policies to address issues of environmental justice and, more generally, for social scientists interested in the determinants of demand for environmental goods.

References

Alesina, Alberto, Batir, Reza, and William Easterly (1999): “Public Goods and Ethnic Divisions,” the *Quarterly Journal of Economics,* 114: 12431284.

Alesina, Alberto and Eliana La Ferrara. “Participation in Heterogeneous Communities,” the *Quarterly Journal of Economics* 3 (2000): 847-904.

Cutler, David, D. Elmendorf, and R. Zeckhauser (1993): “Demographic Characteristics and the Public Bundle,” *Public Finance*, 48: 178-198.

Deacon, Robert and Perry Shapiro (1975): “Private Preference for Collective Goods Revealed Through Voting on Referenda,” *American Economic Review*, 65: 943-955.

Dubin, Jeffrey A., D. Roderick Kiewiet, and Charles Noussair (1992): “Voting on growth control measures: Preferences and strategies,” *Economics and Politics*, 4(2): 191-213.

Easterly, W. and, Levine, R. "Africa's Growth Tragedy: Policies and Ethnic Divisions." *Quarterly Journal of Economics*, 112 (1997): 1203-125.

Fischel, William A. (1979): “Determinants of Voting on Environmental Quality: A Study of a New Hampshire Pulp Mill Referendum,” *Journal of Environmental Economics and Management*, 6: 107-118.

Gates, Gary J., and Jason Ost (2004): *The Gay and Lesbian Atlas*, The Urban Institute Press, Washington, D.C.

Institute for Governmental Studies at the University of California Berkeley, http://www.igs.berkeley.edu

Kahn, Matthew (2002): “Demographic Change and the Demand for Environmental Regulation,” *Journal of Policy Analysis and Management*, 21(1): 45-62.

Kahn, Matthew E., and John G. Matsusaka (1997): “Demand for environmental goods: Evidence from voting patterns on California initiatives,” *Journal of Law and Economics*, 40(1): 137-173.

Matsusaka, John G. (1992): “Economics of Direct Legislation,” *Quarterly Journal of Economics*, 107: 541-571.

Okten, Cagla; Osili, Una Okonkwo (2004): “Contributions in Heterogeneous Communities: Evidence from Indonesia,” *Journal of Population Economics*, 17(4): 603-26.

University of California, Hastings College of the Law. “California Ballot Measures Database,”http://library.uchastings.edu/library/Research%20Databases/CA%20Ballot%20Measures/ca_ballot_measures_main.htm

Vigdor, Jacob L. "Community Composition and Collective Action: Analyzing Initial Mail Response to the 2000 Census." *The Review of Economics and Statistics*, February 2004, 86(1): 303-312.

Vigdor, Jacob L. (2002): "Interpreting Ethnic Fragmentation Effects," *Economics Letters*, 75: 271-276.

Table 1: Summary Statistics, Census Tracts

Variable	Description	N	Mean (Standard Dev.)
yes12	Percent votes in favor of proposition 12, parks, clean water, clean air, and coastal protection	6972	.66 (.133)
yes13	Percent votes in favor of proposition 13, drinking water, clean water, watershed and flood protection	6968	.68 (.125)
yes22	Percent votes in favor of proposition 22, limits on marriage	6956	.61 (.150)
yes25	Percent votes in favor of proposition 25, campaign reform	6961	.35 (.059)
POPULATION	Population	7000	4813.1 (2136.5)
DENSITY	People per square meter	7000	.003 (.0036)
HINCOME	Median Household Income ($10,000)	7000	51.41 (24.75)
PCTWHITE	White Population (percent of total)	6995	.60 (.226)
PCTBLACK	African-American Population (percent of total)	6995	.066 (.116)
PCTNATIVE	Native American Population (percent of total)	6995	.009 (.018)
PCTASIAN	Asian and Pacific Population (percent of total)	6995	.11 (.132)
PCTHISPANIC	Hispanic Population (percent of total)	6995	.31 (.257)
RACE	Race Homogeneity Index (five racial groups)	6995	.62 (.171)
PCTPOP65	Population age 65 or above (percent of total)	6995	.15 (.081)
PCTCOLLEGE	Population with college degree (percent of population 25 years old or older)	6995	.26 (.189)
PCTHS	Population with high-school degree or equivalent (percent of population 25 years old or older)	6995	.50 (.193)
PCTHSDROP	High-school dropouts (percent of population 25 years old or older)	6995	.24 (.193)
EDUCATION	Education Homogeneity Index (three education groups)	6995	.36 (.059)
PCTAGRI	Population employed in agriculture and mining (percent of population 16 years old or older)	7000	.023 (.060)
PCTMANUF	Population employed in manufacturing (percent of population 16 years old or older)	7000	.131 (.079)
PCTFIRE	Population employed in finance, insurance, and real estate (percent of population 16 years old or older)	7000	.066 (.039)
PCTREP	Votes for Republican candidates in 2000 primary elections (percent of total)	6989	.496 (.199)
TURNOUT	Voting-age turnout rate: total votes divided by voting-age population	6992	.28 (.132)

Table 2: Base Models

	(1)	(2)	(3)	(4)
	yes12	yes13	yes22	yes25
PCTMANUF	-.0081	-.0005	.2155***	-.1013***
	(.0119)	(.0118)	(.0285)	(.0162)
PCTAGRI	-.1058***	-.0811***	.0869***	-.1010***
	(.0173)	(.0170)	(.0232)	(.0149)
PCTFIRE	.0644***	.0660***	.1259***	-.0174
	(.0229)	(.0248)	(.0419)	(.0270)
PCTPOP65	.0559***	.1150***	.1789***	-.0020
	(.0092)	(.0088)	(.0110)	(.0087)
PCTBLACK	-.0273***	.0148**	.3246***	-.1366***
	(.0067)	(.0067)	(.0109)	(.0080)
PCTNATIVE	-.0511	.0253	.2849*	.0048
	(.0345)	(.0545)	(.1704)	(.0885)
PCTASIAN	-.0100	.0140**	.2672***	.0404***
	(.0061)	(.0058)	(.0098)	(.0059)
PCTHISPANIC	.0444***	.0595***	.1814***	-.0242***
	(.0070)	(.0070)	(.0135)	(.0090)
PCTCOLLEGE	.2499***	.2286***	-.4572***	.1374***
	(.0131)	(.0135)	(.0186)	(.0133)
PCTHS	-.0075	-.0128	-.0096	.0377***
	(.0137)	(.0137)	(.0155)	(.0121)
PCTREP	-.5620***	-.5072***	.5880***	-.0993***
	(.0081)	(.0083)	(.0141)	(.0095)
HINCOME	-.0160***	-.0130***	.0251***	-.0108***
	(.0015)	(.0016)	(.0025)	(.0018)
HINCOME* HINCOME	.0008***	.0007***	-.0008***	.0006***
	(.0001)	(.0001)	(.0001)	(.0001)
DENSITY	3.9705***	3.5422***	-1.2021*	-1.7719***
	(.4777)	(.6116)	(.6236)	(.3944)
DENSITY* DENSITY	-79.6473***	-25.2438	18.1746	61.1955***
	(15.6464)	(27.7228)	(2.4517)	(12.3688)
TURNOUT	-.0750***	-.0711***	.0662***	-.1736***
	(.0142)	(.0139)	(.0203)	(.0129)
Constant	.9303***	.9154***	.1694***	.4785***
	(.0110)	(.0136)	(.0195)	(.0137)
Observations	6965	6961	6949	6954
LL	13512.96	13004.27	10477.2	12754.0
BIC	-26371.13	-25353.78	-20317.47	-24862.17
R-squared	.9259	.9034	.8617	.4963

Robust Standard errors in parentheses. All models are weighted by the population of each tract and include county dummies. * significant at 10%; ** significant at 5%; *** significant at 1%.

Table 3: Effects of Race Homogeneity

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	Yes12	Yes12	yes13	yes13	yes22	yes22	yes25	yes25
PCTBLACK	-.0085	.1004***	.0281***	.1455***	.3697***	.4049***	-.1443***	-.0975***
	(.0080)	(.0195)	(.0081)	(.0202)	(.0121)	(.0340)	(.0090)	(.0235)
PCTNATIVE	-.0281	.0222	.0415	.0322	.3399**	-.0325	-.0046	.2014***
	(.0350)	(.0538)	(.0541)	(.0677)	(.1579)	(.0746)	(.0869)	(.0576)
PCTASIAN	.0126*	.1161***	.0301***	.1253***	.3214***	.3944***	.0311***	.0881***
	(.0073)	(.0211)	(.0072)	(.0211)	(.0127)	(.0367)	(.0078)	(.0253)
PCTHISPANIC	.0448***	.0471***	.0598***	.0506***	.1824***	.0611**	-.0244***	-.0055
	(.0071)	(.0156)	(.0071)	(.0155)	(.0136)	(.0239)	(.0090)	(.0157)
Race Homogeneity Index	.0327***		.0232***		.0783***		-.0135**	
	(.0060)		(.0061)		(.0081)		(.0055)	
White share squared		.0741***		.0605***		.0540***		.0157
		(.0101)		(.0104)		(.0183)		(.0123)
Black share squared		-.0703***		-.0988***		.0090		-.0410*
		(.0185)		(.0188)		(.0296)		(.0229)
Native share squared		.0416		.1816		1.0466***		-.4713***
		(.0788)		(.1124)		(.1572)		(.1195)
Asian share squared		-.0689***		-.0699***		-.0859***		-.0584**
		(.0246)		(.0238)		(.0323)		(.0250)
Hispanic share squared		.0657***		.0677***		.2003***		-.0113
		(.0161)		(.0163)		(.0166)		(.0131)
Constant	.9015***	.8610***	.7458***	.7007***	.0827***	.1038***	.4935***	.4595***
	(.0127)	(.0154)	(.0128)	(.0155)	(.0238)	(.0306)	(.0169)	(.0195)
Observations	6965	6965	6961	6961	6949	6949	6954	6954
LL	13540.07	13571.5	13016.15	13050.65	10543.19	10618.41	12757.72	12771.78
BIC	-26416.49	-26461.66	-25368.7	-25411.15	-20440.61	-20555.65	-24860.76	-24844.64
R-squared	.9265	.9271	.9037	.9047	.8643	.8672	.4968	.4988

Robust Standard errors in parentheses. All models are weighted by the population of each tract and include county dummies. * significant at 10%; ** significant at 5%; *** significant at 1%.

Table 4: Effects of Education Homogeneity

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	yes12	yes12	yes13	yes13	yes22	yes22	yes25	yes25
PCTCOLLEGE	.2547***	.3762***	.2331***	.3623***	-.4444***	-.3657***	.1371***	.0455
	(.0139)	(.0339)	(.0142)	(.0346)	(.0180)	(.0364)	(.0135)	(.0288)
PCTHS	-.0008	.0987*	-.0066	.1851***	.0084	.9398***	.0373***	-.2909***
	(.0145)	(.0560)	(.0144)	(.0603)	(.0152)	(.0937)	(.0124)	(.0713)
Education Homogeneity Index	.0644***		.0599***		.1703***		-.0041	
	(.0195)		(.0188)		(.0154)		(.0138)	
College share squared		-.0171		-.0002		.3894***		-.0220
		(.0235)		(.0256)		(.0420)		(.0315)
High-school share squared		.0361		-.0318		-.4684***		.1984***
		(.0552)		(.0550)		(.0675)		(.0557)
High-school dropout share squared		.1632***		.1915***		.5010***		-.1566***
		(.0298)		(.0306)		(.0385)		(.0290)
Constant	.8909***	.8147***	.8788***	.7660***	.0530**	-.3067***	.4813***	.6267***
	(.0176)	(.0262)	(.0189)	(.0293)	(.0217)	(.0387)	(.0170)	(.0288)
Observations	6965	6965	6961	6961	6949	6949	6954	6954
LL	13542.9	13573.79	13026.67	13059.37	10565.73	10648.06	12754.1	12781.57
BIC	-26431.0	-26466.22	-25389.73	-25446.28	-20485.68	-20632.64	-24862.36	-24899.62
R-squared	.9265	.9272	.9040	.9049	.8657	.8684	.4963	.50

Robust Standard errors in parentheses. All models are weighted by the population of each tract. All models include county dummies and additional controls in Table 2. * significant at 10%; ** significant at 5%; *** significant at 1%.

Table 5: Split Sample, Effects of Race Homogeneity on Initiative 12

	(1)	(2)	(3)	(4)	(5)	(6)
	25 Percentile (least homogenous)	IQR	75 Percentile (most homogenous)	25 Percentile (least homogenous)	IQR	75 Percentile (most homogenous)
	yes12	yes12	yes12	yes12	yes12	yes12
PCTBLACK	-.0289*	-.0066	-.0164	.0655	.1054***	.0628
	(.0169)	(.0095)	(.0148)	(.0550)	(.0293)	(.0696)
PCTNATIVE	.0657	-.1033***	-.1821	.4279**	-.0395	.0236
	(.1022)	(.0351)	(.1181)	(.2138)	(.0652)	(.2446)
PCTASIAN	.0150	.0293**	.0007	.1440**	.1116***	.1115**
	(.0157)	(.0122)	(.0261)	(.0587)	(.0279)	(.0501)
PCTHISPANIC	-.0062	.0665***	.0685***	-.0493	.0309	-.0138
	(.0197)	(.0086)	(.0108)	(.0527)	(.0216)	(.0311)
Race Homogeneity Index	.0373	.0343***	.0239			
	(.0325)	(.0099)	(.0152)			
White share squared				.0515	.0611***	-.0224
				(.0543)	(.0156)	(.0309)
Black share squared				-.0454	-.0892***	-.1138**
				(.0479)	(.0274)	(.0568)
Native share squared				-.7638	-.0489	-2.1512
				(.5725)	(.0739)	(6.3725)
Asian share squared				-.1097**	-.0454	-.1742***
				(.0463)	(.0349)	(.0538)
Hispanic share squared				.1635**	.0999***	.0805***
				(.0736)	(.0222)	(.0267)
Constant	.7726***	.7137***	.7383***	.6942***	.9141***	.7775***
	(.0528)	(.0190)	(.0220)	(.0511)	(.0188)	(.0341)
Observations	1746	3483	1736	1746	3483	1736
R-squared	.8727	.9323	.9571	.8569	.9326	.9505

Robust Standard errors in parentheses. All models are weighted by the population of each tract. All models include county dummies and additional controls in Table 2. * significant at 10%; ** significant at 5%; *** significant at 1%.

Table 6: Split Sample, Effects of Race Homogeneity on Initiative 13

	(1)	(2)	(3)	(4)	(5)	(6)
	25 Percentile (least homogenous)	IQR	75 Percentile (most homogenous)	25 Percentile (least homogenous)	IQR	75 Percentile (most homogenous)
	yes13	yes13	yes13	yes13	yes13	yes13
PCTBLACK	-.0013	.0342***	.0160	.0880	.1858***	.1985***
	(.0169)	(.0098)	(.0162)	(.0553)	(.0306)	(.0583)
PCTNATIVE	.2461*	-.0609	-.2254*	.4125*	.0433	.1830
	(.1284)	(.0389)	(.1269)	(.2145)	(.1140)	(.2446)
PCTASIAN	.0340**	.0437***	.0286	.1352**	.1415***	.1684***
	(.0154)	(.0108)	(.0263)	(.0576)	(.0282)	(.0599)
PCTHISPANIC	-.0037	.0810***	.0863***	-.0377	.0500**	.0647*
	(.0195)	(.0083)	(.0117)	(.0507)	(.0225)	(.0346)
Race Homogeneity Index	.0379	.0222**	.0094			
	(.0300)	(.0097)	(.0158)			
White share squared				.0460	.0577***	.0135
				(.0531)	(.0164)	(.0243)
Black share squared				-.0660	-.1499***	-.1762***
				(.0502)	(.0287)	(.0567)
Native share squared				.2595	-.0334	-6.3618
				(.5649)	(.1201)	(6.1147)
Asian share squared				-.0770*	-.0629*	-.1656***
				(.0442)	(.0324)	(.0626)
Hispanic share squared				.1495**	.0866***	.0552*
				(.0716)	(.0229)	(.0291)
Constant	.6663***	.6658***	.6508***	.7163***	.8590***	.7877***
	(.0609)	(.0186)	(.0215)	(.0520)	(.0240)	(.0298)
Observations	1746	3487	1728	1746	3487	1728
R-squared	.8540	.9179	.9260	.8370	.9159	.9224

Robust Standard errors in parentheses. All models are weighted by the population of each tract. All models include county dummies and additional controls in Table 2. * significant at 10%; ** significant at 5%; *** significant at 1%.

Table 7: Split Sample, Effects of Education Homogeneity on Initiative 12

	(1)	(2)	(3)	(4)	(5)	(6)
	25 Percentile (least homogenous)	IQR	75 Percentile (most homogenous)	25 Percentile (least homogenous)	IQR	75 Percentile (most homogenous)
	yes12	yes12	yes12	yes12	yes12	yes12
PCTCOLLEGE	.1749***	.2757***	.2937***	.1522**	.3221***	.5588***
	(.0200)	(.0156)	(.0322)	(.0729)	(.0333)	(.0859)
PCTHS	-.0787***	.0134	.0185	.0811	.2321***	.4045***
	(.0222)	(.0144)	(.0299)	(.1709)	(.0873)	(.1409)
Education Homogeneity Index	-.0136	.1745***	.1023**			
	(.0448)	(.0321)	(.0412)			
College share squared				.0221	.1823***	.0214
				(.0774)	(.0468)	(.0587)
High-school share squared				-.1876	.0261	-.0515
				(.1765)	(.0684)	(.0741)
High-school dropout share squared				-.0181	.2624***	.3946***
				(.0763)	(.0464)	(.0959)
Constant	.7984***	.8242***	.7277***	.7643***	.7327***	.5158***
	(.0258)	(.0180)	(.0396)	(.0603)	(.0385)	(.0799)
Observations	1743	3487	1735	1743	3487	1735
R-squared	.9114	.9251	.9462	.9114	.9253	.9483

Robust Standard errors in parentheses. All models are weighted by the population of each tract. All models include county dummies and additional controls in Table 2. * significant at 10%; ** significant at 5%; *** significant at 1%.

Table 8: Split Sample, Effects of Education Homogeneity on Initiative 13

	(1)	(2)	(3)	(4)	(5)	(6)
	25 Percentile (least homogenous)	IQR	75 Percentile (most homogenous)	25 Percentile (least homogenous)	IQR	75 Percentile (most homogenous)
	yes13	yes13	yes13	yes13	yes13	yes13
PCTCOLLEGE	.1222***	.2425***	.2595***	.0863	.2385***	.5207***
	(.0188)	(.0172)	(.0319)	(.0743)	(.0371)	(.0831)
PCTHS	-.1013***	-.0022	-.0044	.1138	.2301**	.2379
	(.0227)	(.0154)	(.0301)	(.1663)	(.0948)	(.1640)
Education Homogeneity Index	-.0266	.1140***	.1275***			
	(.0457)	(.0351)	(.0466)			
College share squared				.0272	.1746***	-.0091
				(.0761)	(.0531)	(.0731)
High-school share squared				-.2629	-.0641	.0577
				(.1746)	(.0773)	(.0919)
High-school dropout share squared				-.0371	.1713***	.3612***
				(.0797)	(.0477)	(.1035)
Constant	.8092***	.8320***	.7042***	.7656***	.7512***	.5519***
	(.0246)	(.0201)	(.0445)	(.0592)	(.0398)	(.0916)
Observations	1745	3483	1736	1745	3483	1736
R-squared	.8955	.90	.9218	.8956	.9010	.9244

Robust Standard errors in parentheses. All models are weighted by the population of each tract. All models include county dummies and additional controls in Table 2. * significant at 10%; ** significant at 5%; *** significant at 1%.