UNIVERSAL VOUCHERS AND RACIAL AND ETHNIC SEGREGATION

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Abstract—We use data on vote outcomes from a universal voucher initiative to examine whether white households with children in public schools will use vouchers to leave predominantly nonwhite schools, thereby contributing to more racially and ethnically segregated schools. We find that white households are more likely to support vouchers when their children attend schools with larger concentrations of nonwhite schoolchildren, an effect that is absent for nonwhite households and households without children. This result may be driven less by race or ethnicity and more by other characteristics, such as student performance, that are correlated with race or ethnicity.

I. Introduction

In the debate over school choice, one of the chief concerns about universal vouchers is that they may lead to more racially and ethnically segregated schools. Since Brown v. Board of Education, school desegregation efforts have attempted to ensure equal educational opportunities for all students through more integrated schools. In addition, recent empirical evidence suggests that the racial composition of a school is a large determinant of the black-white gap in student performance, with black students performing substantially better when they are in racially mixed environments (Card & Rothstein, 2007; Hanushek, Kain, & Rivkin, 2002; Lutz, 2005). These benefits could be diminished if white households are more inclined to use vouchers to flee from the public sector, creating more segregated schools.

Because no state has yet adopted a comprehensive voucher, we cannot directly predict what the effects of such a policy are likely to be. Instead, existing evidence about the impact of vouchers on racial or ethnic segregation has come from studies that examine school segregation within the context of existing forms of school choice (private schools, charter schools, or limited voucher programs). In general, these studies have produced conflicting evidence on whether expanded school choice will lead to more segregated schools. For example, Conlon and Kimenyi (1991), Lankford, Lee, and Wyckoff (1995), Lankford and Wyckoff (2000), and Fairlie and Resch (2002) find that higher concentrations of black or nonwhite students in area public schools increase the likelihood that white households will enroll their children in private schools, apparent evidence that vouchers will increase racial segregation. In contrast, Campbell, West and Peterson (2005) and Buddin, Cordes, and Kirby (1998), find no evidence of white flight, suggesting that perhaps the concerns of voucher opponents are overstated.

In this paper, we present new evidence on whether universal vouchers will lead to more racially and ethnically segregated schools. Our conceptual framework highlights the difference between the causal effect of school vouchers on racial/ethnic segregation and what we refer to as the policy effect. This policy effect allows for the fact that the racial and ethnic composition of schools is likely correlated with other important aspects of school quality and that households may have already systematically sorted into their current school. Thus, the policy effect provides an indication of the additional sorting, given the choices that households have already made, that we should expect to occur under a comprehensive voucher system. To identify the policy effect, we estimate the likelihood that white households with children currently in public schools will use vouchers to opt out of schools with higher concentrations of nonwhite students.

Unlike previous studies of school segregation that have generally extrapolated from data under existing forms of school choice (for example, private schools or limited voucher programs), we use votes on a statewide universal voucher initiative in California as a proxy for voter intention to use the voucher. Assuming that support for the initiative reflects a desire to take advantage of the voucher, voting patterns can reveal which schools are most likely to lose students if a comprehensive voucher is adopted. One distinct advantage of using these data is that to the extent that voters are nonmyopic, these patterns will take account of the general policy effects (changes in public school quality or the availability of private school options, for example) of a universal voucher.2 That is, a universal voucher, available to all families on a statewide basis, would fundamentally change the institutional structure of school finance and could affect household decisions in ways that limited, targeted programs would not. Rather than extrapolating from existing choices (that might have been very different under an entirely new system), our data may better represent the school choice decisions that households would make under a universal voucher.

We find that a universal voucher is likely to increase racial and ethnic segregation across public schools. White households with children currently in public schools are increasingly supportive of the voucher if their children attend schools that have higher concentrations of nonwhite students, an effect that is absent for nonwhite households

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2 Naturally voters may not be able to foresee all effects of a universal voucher system, but this type of analysis provides a useful alternative to studies that examine the effect of small voucher programs, which likely caused little or no supply response from the private school system. See Epple, Romer, and Sieg (2001) and Calabrese et al. (2006) for a discussion of whether voters are myopic.
with children in public schools and households without school-age children. In subsequent analysis, we find that our results appear to be driven less by race or ethnicity and more by other student characteristics that are correlated with race or ethnicity. For example, consistent with Betts and Fairlie (2003), we find evidence that voucher support among white households increases with the share of nonwhites, particularly Hispanics, who are limited English proficient.

II. Conceptual Framework

The primary motivation for this paper is the question of whether a universal school voucher system will lead to increased racial or ethnic segregation. We focus on the additional stratification that vouchers may cause relative to the level of segregation that already exists under the current school system. Specifically, we differentiate between the causal effect of race on the residential and school choice decisions of white households and what we refer to as the policy effect. The causal effect captures whether school racial/ethnic composition causes white households to opt out of public schools, all else equal. In order to obtain causal estimates, the analyst must be concerned with potential bias from two types of omitted variables: unobserved household preferences that may be correlated with the composition of the school where the household currently resides and unobserved attributes that may be correlated with school composition (such as school quality). Once these biases are removed, the causal effect measures the impact of racial/ethnic composition if white households were randomly assigned to schools that are identical except that they have different concentrations of nonwhite students. In contrast, the policy effect measures the impact of nonwhite concentration on household choices, allowing for the fact that households may have already sorted across neighborhood schools based on their preferences for racial or ethnic homogeneity and allowing for the fact that a school’s composition is likely correlated with other school attributes.

While the causal effect is a legitimate concern in its own right, we argue that it is clearly the policy effect that is most important to policymakers concerned about the stratification effects of adopting a universal voucher. To illustrate that point, consider a behavior that is consistent with vouchers increasing racial and ethnic segregation beyond current levels: white households with children currently in relatively integrated public schools use the voucher to leave the public sector at higher rates than nonwhite households with children in those schools. From the perspective of whether vouchers are likely to lead to more segregated schools, it does not matter why white households are opting out; it only matters that they do. Whether whites are fleeing from nonwhite students or from bad schools, high crime, immigrants, or something else does not really matter. If these are all highly correlated, the end result will still be that whites flee schools with observably more nonwhite students, and universal vouchers will still increase the concentration of nonwhite students in those schools. On the other hand, it does matter that households have already chosen their residential location and school, as parents that are relatively satisfied with their current choice will be less likely to use the voucher to move. To use an extreme example, if there were costless mobility and enough schools such that every household could send their child to a school with a bundle of attributes that exactly matched their preferences (perfect Tiebout sorting), then the adoption of vouchers would have no effect on the behavior of white households, regardless of the racial composition of schools and even if the original sorting had been driven entirely by racial preferences.

Thus, whether vouchers will increase racial and ethnic segregation beyond current levels depends on the relative impact of school racial/ethnic composition (and all variables correlated with racial/ethnic composition) on household choices, allowing for for Tiebout sorting. Given the degree to which households have already sorted and the strong correlation between school nonwhite concentration and many aspects of school quality, the distinction between the causal and policy effects is not a trivial one. For example, studies that control for sorting (Lankford et al., 1995; Fairlie & Resch, 2002) typically find that vouchers increase segregation. These studies, however, most likely overestimate the policy effect because they have controlled for differences in preferences that led white households in schools with substantial nonwhite populations to choose those schools in the first place. Similarly, studies that control for correlated school attributes (Campbell et al., 2005; Buddin et al., 1998) find at most a weak relationship between vouchers and segregation. These studies, however, most likely underestimate the policy effect because they have conditioned away factors that may cause whites to leave schools with predominantly nonwhite populations. The substantially different implications from these two types of studies suggest that the existence and magnitude of the policy effect of a voucher system on school segregation are still undetermined.

A. Estimating the Policy Effect

To develop the empirical implications of this difference between the causal and policy effects for the public school system, we present a simple model of voucher use that

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3 As we discuss in detail later in the paper, for universal vouchers to increase racial segregation systemwide (in public and private schools), it is also necessary that white households opt into private schools with higher concentrations of nonwhite students at a lower rate than do nonwhite households.

4 For example, Downes and Zabel (2002) find that households appear to use easily observable characteristics such as race when comparing public school quality among schools. In addition, Clapp, Nanda, and Ross (2008) find that the conditional correlation between housing prices and test scores is eliminated after controlling for the demographic composition of students. These results suggest that perceived or actual school quality is higher in schools and school districts that contain lower concentrations of nonwhite students.
allows for a correlation between preferences and current school racial/ethnic composition, and only conditions on the racial/ethnic composition of the school. Thus, our simple model allows all relevant school attributes to operate through racial/ethnic composition. Let $Y_{is}$ denote the likelihood that individual $i$ who is a member of group $k$ ($k = \text{white, nonwhite}$) in school $s$ uses the voucher. $Y_{is}$ is assumed to be a linear function of the nonwhite share in the school, $M_s$, the individual’s preferences for racial homogeneity, $P_{is}$, and an idiosyncratic error term, $\epsilon_{is}$:

$$Y_{is} = \beta^e_i + \beta^M_i M_s + P_{is} + \epsilon_{is}. \quad (1)$$

$P_{is}$ may be correlated with $M_s$ to the extent that households have already sorted based on race or ethnicity; that is, households may have already chosen their current school based in part on their preferences for racial or ethnic homogeneity. $\beta^M_i$ is the direct effect of nonwhite concentration on voucher use by group $k$, controlling for these preferences.

The policy effect of a change in nonwhite concentration on voucher use by group $k$ is thus defined as

$$\beta^k_{1, \text{pol}} = \text{Mean}^k \left( \frac{dY_{is}}{dM_s} \right) = \beta^k + \frac{dE[P_{is}|k,s]}{dM_s}. \quad (2)$$

Let us assume that as a result of sorting, the preferences of an individual in school $s$ can be described as a linear function of school nonwhite share and an idiosyncratic error, $\mu_{is}$:

$$P_{is} = \alpha^k + \alpha^M_i M_s + \mu_{is}. \quad (3)$$

The average group preferences in a neighborhood are described by

$$E[P_{is}|k,s] = \alpha^k + \alpha^M_i M_s + E[\mu_{is}|k,s] = \alpha^k + \alpha^M_i M_s, \quad (4)$$

which implies that $\frac{dE[P_{is}|k,s]}{dM_s} = \alpha^k$. We can use equations (2) through (4) to express the policy effect of a change in nonwhite concentration on voucher use:

$$\beta^k_{1, \text{pol}} = \beta^k + \alpha^k. \quad (5)$$

Note that if white households with strong preferences for racial/ethnic homogeneity have already sorted into more homogeneous schools, we would expect $\alpha^w > 0$ (where $w$ denotes white households) to be less than 0. Thus, equation (5) implies that for white households, residential sorting will mitigate the policy impact of nonwhite concentration on voucher use relative to the direct effect alone.

For estimation, we cannot observe preferences, so the expected value of the ordinary least squares estimator of $\beta^k$ can be derived from equation (1) using standard theoretical results on omitted variables with one observable regressor, $M_s$, and simplified using equation (3). Specifically,

$$E \left[ \beta^k_{1, \text{OLS}} | M_s, k \right] = \beta^k + \left( \frac{E[P_{is}|M_s|k]}{E[(M_s^2)^{1/2}|k]} \right) = \beta^k + \alpha^k, \quad (6)$$

which equals the policy effect derived in equation (5). Thus, with data on voucher use and school racial/ethnic composition, we can uncover the policy effect by estimating a very simple model with OLS.

Note that vouchers will increase the racial and ethnic segregation of public schools only if white households use the voucher to leave the public sector at higher rates than nonwhite households. If white and nonwhite families are equally likely to use vouchers to opt out of the public sector when their children attend schools with high concentrations of nonwhite students, then these schools may lose a large number of students, but there may be little effect on the overall nonwhite shares in public schools. Thus, it is the comparison of the OLS estimates for white and nonwhite households that will indicate the potential impact of a universal voucher program on racial and ethnic segregation in public schools.

**B. Will Vouchers Increase Racial Segregation Systemwide?**

Beyond the question of whether a universal voucher system would increase racial and ethnic segregation in public schools is the more general question of whether such a system would lead to a systemwide (public and private school) increase in segregation. If we had data on actual voucher take-up rates and schools chosen in a universal voucher system, we could directly capture any systemwide changes in segregation by examining the level of segregation before and after implementation of the system. Unfortunately, such data do not exist; although several states have considered legislation to establish universal vouchers, none has yet adopted such a policy.\(^5\) It is therefore worth discussing the circumstances under which our model, which focuses on the differential likelihood of voucher take-up among whites and nonwhites at predominantly nonwhite public schools, also identifies higher levels of systemwide segregation. For this discussion, we will use 1 minus the sum of the exposure of white students to nonwhite students and nonwhite students to white students as our measure of school segregation. Specifically:

$$E = 1 - \sum_{\forall s} \frac{(N_{ws}/N_w)M_s + (N_{ms}/N_m)(1 - M_s)}{N_s} = 1 - \sum_{\forall s} \frac{\lambda_{ws}M_s + \lambda_{ms}(1 - M_s)}{N_s}, \quad (7)$$

where the first term in the summation captures white exposure to nonwhites by multiplying the fraction of all white students in the system ($\lambda_{ws} = N_{ws}/N_w$) who attend school $s$ times the share nonwhite ($M_s$) at school $s$ and the second

\(^5\) It should be noted that other countries have adopted fairly comprehensive voucher programs. Hsieh and Urquiola (2006) find that Chile’s voucher program led to increased sorting as a disproportionate number of students from high- and middle-income households opted out of the public sector. Ladd and Fiske (2001) find that the introduction of a quasi-universal voucher system in New Zealand led to increased racial segregation.
term captures nonwhite exposure to whites in a similar manner.6

The effect of a policy change, Δ, on school segregation may be expressed as

\[
\frac{dE}{d\Delta} = \sum_{g} \left[ \lambda_{ms} \frac{\partial M_s}{\partial \Delta} - \lambda_{ws} \frac{\partial M_s}{\partial \Delta} \right]
\]

where

\[
\left[ \frac{\partial M_s}{\partial \Delta} \left( \frac{1}{1 - M_s} \right) \right]_{\lambda_{ms}, \lambda_{ws}}
\]

Focusing on the first term in equation (8), note that λ_{ws} and λ_{ms} can respectively be written as λ_s(1 - M_s)/(1 - M) and λ_s/M, where M is the share nonwhite in the entire system and λ_s is the share of total students at school s. The first term in equation (8) can therefore be expressed as

\[
\frac{\partial E}{\partial \Delta}_{\lambda_{ms}, \lambda_{ws}} = \frac{1}{M(1 - M)} \sum_{g} \frac{\partial M_s}{\partial \Delta} \left( \lambda_s(M_s - M) \right)
\]

where

\[
\left[ \frac{\partial M_s}{\partial \Delta} \lambda_s \right]_{\lambda_{ms}, \lambda_{ws}}
\]

It follows from equation (9) that the impact of universal vouchers on systemwide segregation depends solely on the covariance between ∂M_s/∂Δ and λ_sM_s. Specifically, if schools with a larger concentration of nonwhite students experience an increase in nonwhite concentration after implementation of a voucher program, implying Cov[∂M_s/∂Δ, λ_sM_s] > 0, the partial derivative in equation (9) will be positive, and school segregation will increase.7

Note that systemwide, the covariance between ∂M_s/∂Δ and λ_sM_s will depend on two factors: the effect that voucher users have on the racial/ethnic composition of the schools they leave, and the effect that voucher users have on the racial/ethnic composition of schools they choose to attend. Specifically, segregation will increase across the schools' voucher users leave as long as white households opt out of schools with higher concentrations of nonwhite students at a greater rate than do nonwhite households, and segregation will increase across the schools that voucher users choose to attend as long as white households opt into schools with higher concentrations of nonwhite students at a lower rate than do nonwhite households. Under these two circumstances, the partial derivative in equation (9) will be unambiguously positive, implying that segregation increases for the entire population of students. It is notable that this result does not require that nonwhite students self-segregate by choosing schools that have higher nonwhite shares than the ones that they left. It only requires that nonwhite voucher users choose schools with larger nonwhite concentrations at a higher rate than white voucher users.

Turning to the second term in equation (8), which holds M_s fixed for all schools, note that the partial derivatives of λ_{ws} and λ_{ms} must each sum to 0 because the λ's always sum to 1 over all schools. Thus, the second term simplifies to

\[
\frac{\partial E}{\partial \Delta} = \sum_{g} M_s \left[ \frac{\partial \lambda_{ms}}{\partial \Delta} - \frac{\partial \lambda_{ws}}{\partial \Delta} \right] = \text{Cov} \left[ M_s, \frac{\partial \lambda_{ms}}{\partial \Delta} - \frac{\partial \lambda_{ws}}{\partial \Delta} \right].
\]

Similar to the arguments made above, if schools with a larger concentration of nonwhite students experience an increase in nonwhite concentration, then for any school, higher M_s implies positive values of (λ_{ms} - λ_{ws}), and therefore Cov[M_s, ∂λ_{ms}/∂Δ - ∂λ_{ws}/∂Δ] > 0 and equation (10) is unambiguously positive.8 As before, the more general question of whether segregation will increase for all students, including students in private schools, can be signed by the assumption that white voucher users opt into schools with high concentrations of nonwhites at a lower rate than nonwhite voucher users.

Thus, in order to claim that universal vouchers will lead to increased racial and ethnic segregation across all schools, we must determine that white households opt out of schools with higher concentrations of nonwhite students at a greater rate than do nonwhite households, and white households opt into schools with higher concentrations of nonwhite students at a lower rate than do nonwhite households. In the empirical work presented in the next section, we test for the first effect. Unfortunately, our data do not allow us to infer anything about the schools that families would choose under a universal voucher system so we cannot directly test for the second effect. Nevertheless, the existing literature does provide evidence on the choices families have made under similar circumstances and generally supports the assumption that white households are less likely than nonwhite households to choose schools with higher concentrations of nonwhite students. For example, Bifulco and Ladd (2007), Weiher and Tedin (2002), and Hanushek et al. (2007) examine the choices families make in the context of charter school choice. All three of those studies find that parents of charter school users tend to select schools that are racially more similar to their own children than the public schools they leave.9 Similarly, Hastings, Kane, and

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6 The maximum value of an exposure index is always the share of the group to which people are being exposed in the population. Therefore, the sum of the two exposure indices can fall anywhere between 0 and 1.

7 Technically, this result could be reversed by a very strong negative relationship between the relative size of a school λ_s and M_s so that the product λ_sM_s falls with M_s. In practice, this possibility does not arise because larger schools tend to have greater shares of nonwhite students.

8 Note that if all schools were the same size, the change in (λ_{ms} - λ_{ws}) with respect to the policy experiment would be proportional to the change in M_s guaranteeing a positive correlation under our earlier assumption. The change in (λ_{ms} - λ_{ws}), however, depends on the size of the jurisdiction. For circumstances where jurisdiction size may be correlated with racial composition, differences in school size change the magnitude of (λ_{ms} - λ_{ws}) but not the sign of the expression. Accordingly, the derivative of (λ_{ms} - λ_{ws}) is still negative for small values of M_s and positive for large values of M_s, and the resulting sum remains unambiguously positive.

9 Saporito (2003) and Bifulco, Ladd, and Ross (2009) examine the choices made by families in the context of magnet school choice. Similar to the studies on charter school choice, these studies find that white families tend to select schools that are “whiter” than the schools they attempt to leave.
Staiger (2006) examine the choices made by families in the context of intradistrict choice. They conclude that “the average preferred school for each racial group was one in which 70%–80% of the school was their own race.”

With that in mind, we now turn to estimating the policy effect using both survey data on voting preferences and aggregate data on actual vote outcomes from a statewide universal voucher initiative in California. Specifically, we estimate equation (1) using votes for California’s Proposition 38 as a proxy for voucher take-up. Proposition 38 was a statewide ballot initiative in 2000 that would have provided families with a scholarship for every child enrolled in a private school. The scholarship would have been the greatest of three amounts: $4,000, half the national average of public school spending per pupil, or half California’s public school spending per pupil. The initiative placed few conditions on scholarship-redeeming schools and prohibited the state from placing additional conditions on these schools in the future. Because the scholarship would have been made available to all students, including those already enrolled in private schools, Proposition 38 would have created the first universal voucher system in the United States.

III. Analysis Using Survey Data

During August, September, and October 2000, the Public Policy Institute of California (PPIC) surveyed 6,022 potential voters concerning issues related to the November 2000 ballot. The surveys were conducted by telephone, using a random-dialing procedure, and were restricted to people age 18 or older. Baldassare (2000) compares the distributions of various characteristics among survey respondents with the distributions of those characteristics from the 2000 Census. He finds that the survey distributions are quite similar to the census distributions, indicating that the surveys were successful in obtaining a representative sample of California residents.

Our focus is on the voting behavior of households with children in public school. Specifically, we seek to ascertain whether white households with children in public school were more likely than nonwhite households to support the voucher if their child attended a school with a larger concentration of nonwhite students. It seems reasonable to assume that someone who wishes to use the voucher to remove their children from public school would vote for Proposition 38. If we assume that voters are cognizant of the general policy effects that may follow the adoption of a comprehensive voucher (such as changes in public school quality and composition and changes in the number and location of private schools), then we can use those votes to predict which schools are most likely to lose students under the new policy. At the same time, we recognize that voting on Proposition 38 may also have been influenced by factors other than a direct desire to use the voucher. For example, people may have voted for or against the initiative due to their political affiliation (Republican or Democrat) or because they expected it to affect their property values. Because we want to capture voucher take-up, we need to expand equation (1) to control for any variables that might affect voting for reasons other than a direct desire to use the voucher.

Thus, our baseline specification contains nonwhite concentration in a respondent’s local public school and a small set of other variables that are likely to affect voting on the voucher initiative. School nonwhite concentration is measured as 1 minus the fraction of students at a school who are classified as non-Hispanic white. We also include indicator variables for political affiliation, gender, and home ownership status. The political affiliation indicator takes a value of unity if the respondent reported being a registered Republican and is included to account for the fact that school vouchers are a mainstay of conservative political ideology. The gender indicator takes the value of unity if a voter is female and is included to control for systematic gender differences in political ideology and voter support for school vouchers. The home ownership variable captures the fact that vouchers may affect property values. In a system where households must live in a particular neighborhood in order to attend a particular school, it is well established that school quality will be capitalized into housing values (Black, 1999). As Nechyba (2000, 2003) noted, because vouchers decouple the link between residential location and school quality, the introduction of a universal school voucher system would cause property values to rise in areas with low-quality schools and to fall in areas with high-quality schools. This creates incentives for home owners to vote for or against vouchers, depending on where they reside. To account for unobservable differences in political ideology between rural and nonrural voters and the fact that capitalization effects are likely to be weak or nonexistent in rural areas where the supply of housing is relatively elastic, we also include an indicator variable that takes the value of unity if a school is located in a rural area. Finally, we include a set of eight regional fixed effects in our analysis to control for unobservable regional variation in support for school vouchers.

10 Note that 26 respondents did not report their home ownership status. Rather than exclude these respondents from our analysis, we created a dummy variable that takes the value of unity if observations on home ownership status were missing. We then included the home ownership variable (with missing values recoded to 0) and the missing value dummy variable in our analysis.

11 Numerous studies have found that women tend to be more liberal than men and express more liberal opinions on social and economic issues (Shapiro & Mahajan, 1986; Deitch, 1988). Furthermore, Belfield (2003) and Paddock and Paddock (2004) find that women tend to be more ideologically opposed to school vouchers than men.

12 This prediction is supported by Brunner, Sonstelie, and Thayer (2001) and Brunner and Sonstelie (2003), who find that home owners are significantly less likely to support school vouchers if they live in a good school district.

13 See Reback (2005) and Hilber and Mayer (2004) for evidence on capitalization effects in rural areas.

14 Each region consists of contiguous counties; see Betts, Reuben, and Danenberg (2000) for a full description.
A key limitation of the PPIC survey is that it does not identify whether a child is currently enrolled in an elementary, middle, or high school, so we cannot analyze the impact of nonwhite concentration on support for school vouchers separately for households with children in primary or secondary school. As Betts and Fairlie (2003) noted, white households may respond differently to the share of nonwhite students at the primary and secondary levels for a number of reasons. First, because several elementary schools typically feed into one high school, parents have more public school choice at the primary level than at the secondary level. Thus, residential sorting based on preferences for racial/ethnic homogeneity is likely to be more pronounced at the primary level. At the secondary level, white children are likely to experience an increase in their exposure to nonwhite students since high schools integrate students from different neighborhoods and different elementary school attendance zones. As noted in section II, increased residential sorting will mitigate the policy impact of racial and ethnic composition on voucher use relative to the causal effect, implying that the impact of nonwhite concentration on support for school vouchers should be larger at the high school level. Second, the extent to which limited English proficient students are placed in mainstream classes (rather than separate English immersion or bilingual classes) increases significantly at the high school level. According to Rossell (2002), 21% of elementary English learner students in California were in mainstream classrooms in 2000, and nearly 50% of secondary English learners were in mainstream classrooms. Thus, the classroom exposure of native white students to English learners is likely to increase at the high school level. Given that the vast majority of limited English proficient students in California are of Hispanic descent, this provides an additional channel through which nonwhite concentration may differentially affect voucher take-up rates at the primary and secondary levels.\footnote{On the other hand, secondary schools are more likely than primary schools to track students into separate classrooms based on ability. Such tracking may mitigate some of the concerns facing white parents if tracking tends to reduce the exposure of white students to students with limited English proficiency or low student performance. See Figlio and Page (2002) for a discussion of the issues related to tracking and evidence on the impact of tracking on student performance.} Finally, the fraction of school-age children who attend private schools is higher at the primary level than at the secondary level. According to data compiled by the California Department of Education, 11% of all elementary students (grades K–6) attended private school in 2000 while 10% of middle school students (grades 7–8) and 8% of high school students (grades 9–12) attended private school. Thus, the propensity of families to use vouchers to send their children to private schools may differ systematically at the primary and secondary levels.

A second and related limitation of the survey data is that they do not identify the public school attended by each respondent’s child. To address this limitation, we used data from the 2000 Census geographic files to match the postal code of each respondent to the closest public elementary school and the closest public high school. Specifically, we first matched the postal code of each respondent to a corresponding school district. We then matched the centroid of the postal code to the closest public elementary and high school within that school district. In cases where a postal code crossed district boundaries, we matched the code to the closest school (elementary and high school) in each of the districts associated with the code. We then weighted the characteristics of each school (fraction nonwhite, and so on) by the fraction of the postal code’s population residing in that school district. Using this procedure, we were able to assign detailed public school characteristics to each respondent in our sample.\footnote{All school-level data are for 1999–2000 and were obtained from the California Department of Education.}

Because we cannot exactly identify the school attended by a respondent’s child, our measures of nonwhite concentration likely suffer from two types of measurement error. The first results from our inability to precisely match respondents to school districts where postal codes overlap school district boundaries. In these cases, our measure of nonwhite concentration is a weighted average, not the actual value of the school attended by the respondent’s child. Measurement error will tend to bias the parameter estimate on nonwhite concentration toward 0. A second type of measurement error results from our inability to precisely match respondents to school attendance zones. This is likely to be a much greater problem at the primary than at the secondary level, since high schools are much larger than elementary schools and tend to have attendance zones that span much larger geographic areas.\footnote{Comparing actual white exposure to nonwhites with the exposure arising in the survey data where school is imputed, we find that these exposure rates have a correlation of 0.97 at the high school level and only 0.89 at the elementary school level. Details are available on request.}

In light of that fact, we proceed by focusing primarily on the case where nonwhite concentration is measured at the high school level and present only a limited number of results where it is measured at the elementary level. Furthermore, we restrict our analysis to include only respondents who live in a postal code where at least 75% of the population resides within the boundaries of a single school district.\footnote{Consistent with the notion that measurement error tends to bias our parameter estimates toward zero, we find that if we relax this sample restriction and use all observations, the estimated coefficient on nonwhite concentration declines in magnitude but remains statistically significant at the 5% level. Similarly, if we tighten the sample restriction and use only respondents who lived in a postal code where at least 85% or 95% of the population resided within the boundaries of a single district, the estimate coefficient on nonwhite concentration increases in magnitude and remains statistically significant at the 5% level. Details available on request.}

A key limitation of the PPIC survey is that it does not identify whether a child is currently enrolled in an elementary, middle, or high school, so we cannot analyze the impact of nonwhite concentration on support for school vouchers separately for households with children in primary or secondary school. As Betts and Fairlie (2003) noted, white households may respond differently to the share of nonwhite students at the primary and secondary levels for a number of reasons. First, because several elementary schools typically feed into one high school, parents have more public school choice at the primary level than at the secondary level. Thus, residential sorting based on preferences for racial/ethnic homogeneity is likely to be more pronounced at the primary level. At the secondary level, white children are likely to experience an increase in their exposure to nonwhite students since high schools integrate students from different neighborhoods and different elementary school attendance zones. As noted in section II, increased residential sorting will mitigate the policy impact of racial and ethnic composition on voucher use relative to the causal effect, implying that the impact of nonwhite concentration on support for school vouchers should be larger at the high school level. Second, the extent to which limited English proficient students are placed in mainstream classes (rather than separate English immersion or bilingual classes) increases significantly at the high school level. According to Rossell (2002), 21% of elementary English learner students in California were in mainstream classrooms in 2000, and nearly 50% of secondary English learners were in mainstream classrooms. Thus, the classroom exposure of native white students to English learners is likely to increase at the high school level. Given that the vast majority of limited English proficient students in California are of Hispanic descent, this provides an additional channel through which nonwhite concentration may differentially affect voucher take-up rates at the primary and secondary levels.\footnote{On the other hand, secondary schools are more likely than primary schools to track students into separate classrooms based on ability. Such tracking may mitigate some of the concerns facing white parents if tracking tends to reduce the exposure of white students to students with limited English proficiency or low student performance. See Figlio and Page (2002) for a discussion of the issues related to tracking and evidence on the impact of tracking on student performance.} Finally, the fraction of school-age children who attend private schools is higher at the primary level than at the secondary level. According to data compiled by the California Department of Education, 11% of all elementary students (grades K–6) attended private school in 2000 while 10% of middle school students (grades 7–8) and 8% of high school students (grades 9–12) attended private school. Thus, the propensity of families to use vouchers to send their children to private schools may differ systematically at the primary and secondary levels.

A second and related limitation of the survey data is that they do not identify the public school attended by each respondent’s children. To address this limitation, we used data from the 2000 Census geographic files to match the postal code of each respondent to the closest public elementary school and the closest public high school. Specifically, we first matched the postal code of each respondent to a corresponding school district. We then matched the centroid of the postal code to the closest public elementary and high school within that school district. In cases where a postal code crossed district boundaries, we matched the code to the closest school (elementary and high school) in each of the districts associated with the code. We then weighted the characteristics of each school (fraction nonwhite, and so on) by the fraction of the postal code’s population residing in that school district. Using this procedure, we were able to assign detailed public school characteristics to each respondent in our sample.\footnote{All school-level data are for 1999–2000 and were obtained from the California Department of Education.}

Because we cannot exactly identify the school attended by a respondent’s child, our measures of nonwhite concentration likely suffer from two types of measurement error. The first results from our inability to precisely match respondents to school districts where postal codes overlap school district boundaries. In these cases, our measure of nonwhite concentration is a weighted average, not the actual value of the school attended by the respondent’s child. Measurement error will tend to bias the parameter estimate on nonwhite concentration toward 0.

A second type of measurement error results from our inability to precisely match respondents to school attendance zones. This is likely to be a much greater problem at the primary than at the secondary level, since high schools are much larger than elementary schools and tend to have attendance zones that span much larger geographic areas.\footnote{Comparing actual white exposure to nonwhites with the exposure arising in the survey data where school is imputed, we find that these exposure rates have a correlation of 0.97 at the high school level and only 0.89 at the elementary school level. Details are available on request.}

In light of that fact, we proceed by focusing primarily on the case where nonwhite concentration is measured at the high school level and present only a limited number of results where it is measured at the elementary level. Furthermore, we restrict our analysis to include only respondents who live in a postal code where at least 75% of the population resides within the boundaries of a single school district.\footnote{Consistent with the notion that measurement error tends to bias our parameter estimates toward zero, we find that if we relax this sample restriction and use all observations, the estimated coefficient on nonwhite concentration declines in magnitude but remains statistically significant at the 5% level. Similarly, if we tighten the sample restriction and use only respondents who lived in a postal code where at least 85% or 95% of the population resided within the boundaries of a single district, the estimate coefficient on nonwhite concentration increases in magnitude and remains statistically significant at the 5% level. Details available on request.}
households that were eligible for the Children’s Scholarship Fund (CSF). Their results suggest that, relative to whites, African Americans and Hispanics were significantly more likely to apply for CSF vouchers. Second, white children tended to attend schools with substantially lower concentrations of nonwhite students than the schools attended by nonwhite children. That pattern is consistent with the notion that there is already significant racial stratification across California’s public schools.

A. Baseline Results

We examine the relationship between nonwhite concentration and support for school vouchers using a linear probability model. Table 2 reports parameter estimates for the sample of white households with children in public school when nonwhite concentration is measured at the high school level. The standard errors for all analysis with the survey data are clustered at the school district level to account for within-district autocorrelation of the disturbance term. We start with the simplest possible model, estimating equation (1) with nonwhite share alone. As shown in column 1, the coefficient on nonwhite concentration is positive and statistically significant at the 5% level, suggesting that white households with children in public school were more likely to support the voucher if their child attended a school with a higher concentration of nonwhite students. However, voting on the voucher may be influenced by factors other than a direct desire to use the voucher, so we want to control for those factors. Our preferred specification therefore includes a limited set of control variables; those results are reported in column 2. The coefficient on nonwhite concentration is again positive and statistically significant at the 5% level, suggesting that nonwhite households with children in public school were more likely to support the voucher if their child attended a school with a higher concentration of nonwhite students. We eliminated these respondents from the analysis. 20 The summary statistics reported in table 1 are based on the sample that uses high school nonwhite concentration. Summary statistics based on elementary school nonwhite concentration are quite similar to those reported in table 1.

19 Among the 1,636 respondents with children in public school, 182 did not answer the question about how they intended to vote on the voucher initiative, 29 refused to report their race, and an additional 58 respondents reported a postal code that either did not exist or was located in a state other than California. We eliminated these respondents from the analysis.

20 The summary statistics reported in table 1 are based on the sample that uses high school nonwhite concentration. Summary statistics based on elementary school nonwhite concentration are quite similar to those reported in table 1.
to support school vouchers; a 10 percentage point increase in the share of nonwhite students is predicted to increase support for the voucher among white households with children in public school by approximately 2.6 percentage points.

The signs of the coefficients on the control variables listed in column 2 are also generally consistent with expectations. For example, the estimated coefficient on Republican is positive and statistically significant. Similarly, consistent with previous studies that examine support for school vouchers, the estimated coefficient on female is negative and statistically significant at the 10% level. The estimated coefficient on rural is positive and statistically significant at the 10% level, which may reflect the fact that rural white voters tend to be more conservative (along unobservable dimensions not captured by political affiliation) and considerably more likely to support conservative political ideology than nonrural white voters.21 Alternatively, the positive coefficient on rural may reflect the fact that rural home owners are less concerned about the impact of school vouchers on their property values since changes in school quality are unlikely to be capitalized into housing values in rural areas.

Column 3 of table 2 presents results from an expanded specification that adds a standard set of demographic variables found in the literature on public-private school choice (Figlio & Stone, 2001; Lankford & Wyckoff, 2000; Fairlie & Resch, 2002). These include family income and an indicator variable that takes the value of unity if a parent is college educated.22 These variables may affect voting behavior through ideological preferences but are also quite likely to affect desire to use the voucher directly, and we therefore include them as a specification check only. The inclusion of these additional explanatory variables causes the coefficient on the nonwhite concentration variable to decline only slightly.

### B. Nonwhite Households with Children

The results reported in table 2 consistently suggest that white households with children in public school are more likely to support school vouchers if their child attends a school with a higher concentration of nonwhite students. This alone, however, does not necessarily imply that vouchers would lead to more racially and ethnically segregated public schools. In particular, if nonwhite families in similar schools are equally likely to opt out of the public sector, then vouchers may have little effect on the overall racial/ethnic composition of public schools. Thus, we must also examine the voting behavior of nonwhite households with children in public school. We estimate our preferred, baseline specification (from column 2 of table 2) using the sample of nonwhite households and report those results in column 2 of table 3. The coefficient on the nonwhite share variable in column 2 is negative and statistically insignificant. Thus, in contrast to white households with children in public school, nonwhite households appear less likely to support school vouchers if their children attend schools with higher concentrations of nonwhites. Taken together, the results reported in columns 1 and 2 therefore suggest that universal vouchers would lead to more segregated public schools.

### C. Households without School-Age Children

We have assumed that parents who wish to use the voucher would vote for Proposition 38 and have included several variables to control for factors that may affect a household’s vote on the voucher initiative other than a direct desire to use the voucher. Nevertheless, it is still possible that white households residing in areas with high concentrations of nonwhites may be voting for the voucher initiative not because they intend to use the vouchers but because of other unobservable factors. To examine that possibility, we estimated our baseline specification using the sample of white households with no school-age children. Assuming that our results are being driven primarily by a desire among households with children to use school vouchers, we should observe a much stronger reaction to nonwhite concentration among white households with school-age children than among those without school-age children.

Results for the sample of white households with no school-age children are reported in column 3 of table 3. The coefficient on the nonwhite share variable in column 3 is statistically insignificant and quite small in magnitude, giving us increased confidence that our results are capturing a relationship between support for school vouchers and

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21 See McKee (2007) for a discussion of the differences in political ideology between rural white voters and nonrural white voters.

22 In the PPIC survey, respondents were asked to report their family income in terms of six mutually exclusive ranges (for example, income of $40,000 to $59,999). We used the midpoints of these ranges to assign a unique income value to each respondent. Forty-five of the respondents in our sample did not answer the question regarding income. We included missing observation flags for these individuals.
nonwhite concentration that is unique to white households with children in public school.\footnote{We also examined whether our results were being driven by the potential impact of vouchers on housing values. Specifically, we interacted school nonwhite share with the home ownership dummy and reestimated all three models reported in table 3. For white and nonwhite households with children in public school, the home owner interaction term was statistically insignificant and its inclusion had little effect on the results reported in columns 1 and 2 of table 3. In general, the elementary-level results are quite similar to those for high schools. Relative to the high school results, the most notable change is the estimated coefficient on nonwhite concentration for white households with children. That coefficient decreases from 0.258 to 0.139 but remains statistically significant at the 10\% level. This decrease is consistent with the notion that measurement error is likely more of a problem when we measure nonwhite concentration at the elementary level.\footnote{An alternative explanation for why the estimated coefficient on nonwhite concentration is smaller at the elementary level is related to sorting. Recall that because several elementary schools typically feed into one high school, parents have more public school choice at the primary level than at the secondary level. Thus, at the primary school level, families may be better able to sort based on their preferences for racial homogeneity, which could mitigate the policy impact of racial composition on voucher use.}\footnote{The third panel of table 4 reports results when we measure nonwhite concentration using the average nonwhite concentration in the closest three elementary schools to the centroid of a respondent’s postal code. Focusing on the three closest elementary schools rather than simply the closest elementary school may help average out some of the measurement error in the elementary school nonwhite composition variable. Consistent with that notion, the results in panel 3 are quite similar to the high school results in panel 1.\footnote{Our view that measurement error explains the attenuation of our estimates for elementary schools is also consistent with the effect of restricting our sample to postal codes where at least 75\% of the postal code’s population resided within the boundaries of a single school district. If we relax that standard, the estimates decrease with the increase in assignment error, and if we tighten the standard to reduce assignment errors, the estimates increase.}}}

\section{Results for Elementary Schools}

The second panel of table 4 reports results when nonwhite concentration is measured using the racial composition of the closest elementary school. For comparison purposes, the first panel of table 4 reproduces the high school results reported in tables 2 and 3.\footnote{For the sake of brevity, table 4 reports only the coefficients on the nonwhite concentration variable; however, we note that all specifications include all the control variables included in column 2 of table 2.} In general, the elementary-level results are quite similar to those for high schools. Relative to the high school results, the most notable change is the estimated coefficient on nonwhite concentration for white households with children. That coefficient decreases from 0.258 to 0.139 but remains statistically significant at the 10\% level. This decrease is consistent with the notion that measurement error is likely more of a problem when we measure nonwhite concentration at the elementary level.

\begin{table}[h]
\centering
\caption{Estimated Coefficient for Nonwhite Households with Children in Public School and Households without Children}
\begin{tabular}{lccc}
\hline
\multicolumn{1}{c}{Regression Specification} & \multicolumn{1}{c}{White with Children} & \multicolumn{1}{c}{Nonwhite with Children} & \multicolumn{1}{c}{White, No Children} \\
\hline
\textit{Closest high school} & & & \\
Nonwhite Share & 0.258** & -0.092 & 0.001 \\
& (0.086) & (0.088) & (0.045) \\
Observations & 672 & 553 & 2,032 \\
\textit{Closest elementary school} & & & \\
Nonwhite Share & 0.139* & -0.027 & -0.003 \\
& (0.079) & (0.089) & (0.041) \\
Observations & 624 & 528 & 1,900 \\
\textit{Three closest elementary schools} & & & \\
Nonwhite Share & 0.214** & -0.043 & 0.015 \\
& (0.084) & (0.091) & (0.044) \\
Observations & 624 & 528 & 1,900 \\
\hline
\end{tabular}
\end{table}

\subsection{From Whom Are Whites Really Fleeing?}

The results reported thus far consistently suggest that adoption of a universal voucher is likely to increase racial and ethnic stratification, as white households with children currently in public schools appear more likely to use the voucher to leave schools with higher concentrations of nonwhites. This leads to the question of whether whites opt out simply to escape from nonwhites or if whites are fleeing from specific characteristics of nonwhite students. For example, table 1 showed that Hispanics are the largest nonwhite group in California. Many of these Hispanic students are immigrants, or children of immigrants, and what we are interpreting as “white flight” may instead reflect “native flight,” similar to the results of Betts and Fairlie (2003), who find that white families are significantly more likely to enroll their children in private school if they live in areas with high concentrations of immigrant children who speak a language other than English at home.

We begin a closer examination of the characteristics of nonwhites by simply separating nonwhite concentration...
into more specific race/ethnicity groups. Due to the measurement error issues associated with using elementary schools, we focus on the case where nonwhite concentration is measured at the high school level. Results are shown in panel 2 of table 5. For comparison, the baseline results from table 3 are repeated in panel 1. When nonwhite share is broken out, the coefficient on Hispanic share is the only race/ethnicity variable that is statistically different from 0, and it is somewhat larger in magnitude than when we use overall nonwhite share. Given that the shares of blacks and Asians are so much smaller than that of Hispanics, it is not particularly surprising that the effect of Hispanics dominates. For example, given the large Hispanic population, it is likely that white households have been less able to “escape” from Hispanic students through residential sorting.

These results are consistent with native flight, particularly given that a large percentage of Hispanic students in California are limited English proficient. In order to examine this possibility more directly, we separate the nonwhite share variable into the fraction of nonwhite students who are limited English proficient and the fraction who are not. As shown in panel 3 of table 5, we find that white households are significantly more likely to support the voucher as the share of LEP nonwhites rises, but there is no effect on non-LEP nonwhites. When we break down LEP and non-LEP nonwhite shares into specific race/ethnicity categories, we find that our results are driven almost entirely by Hispanic LEP students (panel 4 of table 5). These results are highly consistent with the findings of Betts and Fairlie (2003).

We also note that among nonwhite households, support for the voucher is largely unaffected by the school nonwhite share, whether LEP or not. The exception is in schools with high proportions of LEP Asians, which appears to lead to a large reduction in voucher support. This is primarily driven by a few schools with unusually high percentages of LEP Asians.\(^\text{27}\) The difference between white and nonwhite households with children suggests that under a universal voucher program, the largest changes in racial and ethnic stratification are likely to be seen in areas with relatively high proportions of non–English speaking nonwhites.

A related explanation for our results is that limited English proficiency is an easily observable proxy for low student performance. For example, if schools with high concentrations of nonwhite or LEP students tend to have lower average test scores, the results shown in table 5 may simply be picking up the fact that white families are more likely to support the voucher if their child attends a low-performing school. We test this hypothesis directly by adding California’s Academic Performance Index (API) scores to our model; those results are shown in panel 5 of table 5. For white households with children, the coefficient on the API score is negative and significant at the 5% level, while the coefficient on overall nonwhite share declines sharply and is no longer statistically significant.\(^\text{28}\) For nonwhite households, the addition of the API score leads to a stronger negative effect of nonwhite share, indicating that once we have controlled for student performance, nonwhite households appear to have a preference for remaining in schools with more nonwhite students.

### IV. Analysis Using Aggregate Vote Returns

Our primary analysis relies on polling data collected prior to the actual vote on the voucher initiative and thus represents

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\(^{27}\) The mean percentage of LEP Asians in the nonwhite sample is under 3%. However, in six schools, the reported percentage of LEP Asian students is over 20%. When these schools are dropped from the sample, the coefficient on LEP Asian share drops sharply in magnitude and becomes statistically insignificant.

\(^{28}\) When nonwhite share is broken down into the three race categories, none of the race coefficients are statistically significant.
stated versus revealed preferences for school vouchers. That fact raises the potential concern that our polling data may not accurately represent voting behavior. For example, in the case of school vouchers, a voter’s response to a poll may be heavily influenced by his or her political ideology, while his or her actual vote may be more strongly affected by economic factors, such as the net fiscal cost of the voucher program. Concerns over the stated versus revealed nature of our data may be particularly important given the disparity between the fraction of survey respondents who reported they would support the voucher (46%) and the fraction of voters who actually supported the voucher initiative in November (30%). Ideally, we would like to base our analysis on data from exit polls or actual voting data linked to detailed individual characteristics and residential location information. Unfortunately, such data are not available.

However, we do have block-group-level data on the fraction of voters who supported the initiative and their characteristics, data that provide us with the opportunity to conduct a robustness check of the results obtained with the survey data. To illustrate how we use block-group-level vote tallies, consider dividing the voters within block groups into four mutually exclusive groups: white households with school-age children, nonwhite households with school-age children, white households without school-age children, and nonwhite households without school-age children. Specifically, let \( w \) denote whites, \( nw \) denote nonwhites and \( t \) denote households with no school-age children. We specify four separate equations for the fraction of yes votes—one equation for each of the four groups. By aggregating the linear probability model described in equation (1) up to the block group level, the equation for each group \( k \) in block group \( j \) is

\[
y^k_{js} = \beta^k_0 + \beta^k_1 M_j + P^k_{js} + \epsilon^k_{js},
\]

where \( k = \{w, w', nw, mw'\} \), \( y^k_{js} \) denotes the fraction of yes votes on the voucher initiative for voters located in block group \( j \) and school attendance area \( s \), \( M_j \) is the share of nonwhite students at school \( s \), \( P^k_{js} \) is the mean unobserved preference for racial and ethnic homogeneity for voters of type \( k \) located in block group \( j \) (with an expectation of \( E[P^k_{js} | j, k] \)), and \( \epsilon^k_{js} \) is a mean zero average idiosyncratic error term for voters of type \( k \) located in block group \( j \).

Since the proportion of yes votes cast by all four groups must sum to unity, we can express the total proportion of yes votes at the block group level as

\[
y_{js} = \pi^w_{js} \cdot y^w_{js} + \pi^{nw}_{js} \cdot y^{nw}_{js} + \pi^{w'}_{js} \cdot y^{w'}_{js} + \pi^{nw'}_{js} \cdot y^{nw'}_{js},
\]

where \( \pi^k_{js} \) is the fraction of households in group \( k \). Using equation (11), equation (12) can be expressed as

\[
y_{js} = \pi^w_{js} \cdot y^w_{js} + \pi^{nw}_{js} \cdot y^{nw}_{js} + \pi^{w'}_{js} \cdot y^{w'}_{js} + \pi^{nw'}_{js} \cdot y^{nw'}_{js} + \pi^w_{js} \cdot \lambda^w_{js} + \pi^{nw}_{js} \cdot \lambda^{nw}_{js} + \pi^{w'}_{js} \cdot \lambda^{w'}_{js} + \pi^{nw'}_{js} \cdot \lambda^{nw'}_{js} + \pi^w_{js} \cdot \eta_{js} + \pi^{nw}_{js} \cdot \eta^{nw}_{js} + \pi^{w'}_{js} \cdot \eta^{w'}_{js} + \pi^{nw'}_{js} \cdot \eta^{nw'}_{js},
\]

where \( \eta_{js} = P^w_{js} \cdot \pi^w_{js} + P^{nw}_{js} \cdot \pi^{nw}_{js} + P^{w'}_{js} \cdot \pi^{w'}_{js} + P^{nw'}_{js} \cdot \pi^{nw'}_{js} + \epsilon_{js} \) and \( \epsilon_{js} \) is the cross-group weighted sums of the four group-specific errors terms, \( \epsilon^k_{js} \). Thus, \( \eta_{js} \) contains both an idiosyncratic error term and the group-specific unobservable preference components weighted by the fraction of voters in each group. Equation (13) can be estimated using block-group-level vote returns and the characteristics of voters in each block group. The parameters of primary interest are \( \beta^w_1 \) and \( \beta^{nw}_1 \). Specifically, \( \beta^w_1 \) measures how white households with children respond to nonwhite concentration, while \( \beta^{nw}_1 \) measures how nonwhite households with children respond to nonwhite concentration.

We should note that our analysis using aggregate returns is presented solely to provide additional support for the results reported in section III and does not, by itself, represent strong evidence that nonwhite concentration affects voucher support. By estimating the model of individual voting behavior separately by racial status, we were able to focus on the effect of school nonwhite concentration for white and nonwhite households separately. In an aggregate model of voting behavior, however, the demographic composition of the census block group is used to capture the relationship between race/ethnicity and voting, and the expected values of parameter estimates arising from OLS are quite complex.\(^{29}\)

\[\text{A. Data for Aggregate Analysis}\]

The dependent variable in our analysis of aggregate vote returns is the fraction of voters in a block group who supported the voucher initiative. Block-group-level data on vote outcomes for Proposition 38 were obtained from the Statewide Database, maintained by the Institute of Governmental Studies at the University of California, Berkeley. We use voter registration data from the Statewide Database and data from the Census 2000 Summary File 1 (100% population count file) to construct block-group-level variables that match as closely as possible the explanatory variables we used in our analysis of the PPIC survey data: the fraction of households that are home owners, the fraction of voters who are registered Republicans, and the fraction of the voting-age population (individuals 18 years of age or older) who are female. Consistent with equation (13), we also include the fraction of households that are white with school-age children, the fraction of households that are nonwhite with school-age children, the fraction of households that are white with no school-age children, and the fraction of households that are nonwhite with no school-age children. To account for the fact that families with children in private school would have directly benefited from the voucher, we also include the fraction of K–12 students enrolled in private school. Following our analysis using the survey

\[^{29}\text{For the unobservable, the unobservable in equation (13) contains terms repre-}
\[^{29}\text{senting the average preferences of all groups in the neighborhood, and}
\[^{29}\text{these terms are correlated with both } \pi^k_{js} \text{ and } M_j \text{ due to household}
\[^{29}\text{sorting across neighborhoods and schools.} \]
elementary level, our final sample consists of 38,066 block groups that are matched to 847 public high schools. At the sample consists of 25,715 block groups or pseudo-block groups at the high school level, our final sample consists of 25,715 block groups or pseudo-block groups matching to the closest public high school within the appropriate district. When nonwhite concentration is measured at the high school level, 82% of census blocks are matched to the closest elementary school, 47% of these blocks within a block group match to the exact same school. When census blocks are matched to the closest elementary school, 47% of these blocks within a block group are partitioned according to which school district they belong to.

The coefficients of primary interest are those on the interaction terms between school nonwhite concentration and the fractions of white and nonwhite households with school-age children. When nonwhite concentration is measured at the high school level, the coefficient on the interaction between nonwhite concentration and the fraction of white households with children is positive and statistically significant. The mean and standard deviations of the dependent variable (fraction voting yes) and the standard errors reported in table 6 correspond to the total number of voters in each block group. In addition, the standard errors reported in table 6 are adjusted for clustering of the data at the school level. The coefficients of primary interest are those on the interaction terms between school nonwhite concentration and the fractions of white and nonwhite households with school-age children. When nonwhite concentration is measured at the high school level (column 2), the coefficient on the interaction between nonwhite concentration and the fraction of white households with school-age children is positive and statistically significant.

Note that data on the fraction of students attending private school are available only at the block group level. As a result, we assigned the same values of this variable to all pseudo-block groups located in the same block group.

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### Table 6.—Estimated Coefficients Using Aggregate Vote Returns

<table>
<thead>
<tr>
<th>Regression Specification</th>
<th>(1) Mean (s.d.)</th>
<th>(2) Closest High School</th>
<th>(3) Closest Elementary School</th>
<th>(4) Three Closest Elementary Schools</th>
</tr>
</thead>
<tbody>
<tr>
<td>Nonwhite Share × (White with School-Age Children)</td>
<td>0.06 (0.04)</td>
<td>0.223** (0.043)</td>
<td>0.177** (0.024)</td>
<td>0.197** (0.025)</td>
</tr>
<tr>
<td>Nonwhite Share × (Nonwhite with School-Age Children)</td>
<td>0.09 (0.11)</td>
<td>−0.108** (0.052)</td>
<td>−0.042 (0.029)</td>
<td>−0.108** (0.022)</td>
</tr>
<tr>
<td>Nonwhite Share × (White No School-Age Children)</td>
<td>0.24 (0.15)</td>
<td>0.022* (0.014)</td>
<td>0.021** (0.008)</td>
<td>0.021** (0.007)</td>
</tr>
<tr>
<td>Nonwhite Share × (Nonwhite No School-Age Children)</td>
<td>0.15 (0.16)</td>
<td>0.012 (0.045)</td>
<td>0.003 (0.022)</td>
<td>0.058** (0.013)</td>
</tr>
<tr>
<td>Fraction Home owner</td>
<td>0.65 (0.26)</td>
<td>−0.037** (0.003)</td>
<td>−0.033** (0.002)</td>
<td>−0.034** (0.002)</td>
</tr>
<tr>
<td>Fraction Republican</td>
<td>0.36 (0.15)</td>
<td>0.429** (0.012)</td>
<td>0.411** (0.006)</td>
<td>0.410** (0.006)</td>
</tr>
<tr>
<td>Fraction Female</td>
<td>0.52 (0.03)</td>
<td>−0.125** (0.026)</td>
<td>−0.122** (0.015)</td>
<td>−0.120** (0.015)</td>
</tr>
<tr>
<td>Fraction Private</td>
<td>0.14 (0.15)</td>
<td>0.010** (0.004)</td>
<td>0.014** (0.003)</td>
<td>0.011** (0.003)</td>
</tr>
<tr>
<td>Rural</td>
<td>0.07 (0.16)</td>
<td>0.055** (0.006)</td>
<td>0.018** (0.002)</td>
<td>0.015** (0.002)</td>
</tr>
<tr>
<td>Fraction White with School-Age Children</td>
<td>0.14 (0.09)</td>
<td>−0.082** (0.040)</td>
<td>−0.048** (0.019)</td>
<td>−0.014 (0.014)</td>
</tr>
<tr>
<td>Fraction Nonwhite with School-Age Children</td>
<td>0.13 (0.12)</td>
<td>0.146* (0.080)</td>
<td>0.083** (0.041)</td>
<td>0.183** (0.027)</td>
</tr>
<tr>
<td>Fraction White No School-Age Children</td>
<td>0.50 (0.21)</td>
<td>−0.040 (0.041)</td>
<td>−0.039* (0.021)</td>
<td>0.015 (0.014)</td>
</tr>
<tr>
<td>Observations</td>
<td>25,175</td>
<td>25,715</td>
<td>38,066</td>
<td>38,066</td>
</tr>
<tr>
<td>$R^2$</td>
<td>0.69</td>
<td>0.66</td>
<td>0.66</td>
<td>0.67</td>
</tr>
</tbody>
</table>

Robust, clustered standard errors in parentheses. All specifications include regional fixed effects. *Significant at 10%. **Significant at 5%.

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Note that data on the fraction of students attending private school are available only at the block group level. As a result, we assigned the same values of this variable to all pseudo-block groups located in the same block group.

---

30 When census blocks are matched to the closest high school, 82% of these pseudo-block groups are identical to the block group. That is, all blocks within a block group match to the exact same school. When census blocks are matched to the closest elementary school, 47% of these pseudo-block groups are identical to the block group. Block groups containing more than one pseudo-block group arise primarily because some block groups cross district boundaries. In those cases, census blocks within a block group are partitioned according to which school district they belong to.

31 Note that data on the fraction of students attending private school are available only at the block group level. As a result, we assigned the same values of this variable to all pseudo-block groups located in the same block group.

32 Since the summary statistics for the elementary and high schools samples are nearly identical, table 6 reports only summary statistics based on the high school sample. Not reported in table 6 are the means and standard deviations of the dependent variable (fraction voting yes) and the share of nonwhite students in a school. The mean and standard deviation of those variables are 0.29 (0.07) and 0.55 (0.27), respectively.

33 Note that we exclude the level variable (noninteraction variable) for the fraction of nonwhite households with no children in the analysis. Thus, the coefficients on the fraction of white households with children, the fraction of nonwhite households with children, and the fraction of white households with no children are all relative to this omitted group.
significant at the 5% level, indicating that these households are more likely to support the voucher if their child attends a school with a larger concentration of nonwhite students. Furthermore, the point estimate on that interaction is 0.223, quite close to the estimate obtained using the survey data of 0.258. The point estimate on nonwhite concentration for nonwhite households with school-age children (−0.108) is also quite close to the estimate obtained in the survey data (−0.092). Finally, the coefficient on the interaction term between school nonwhite concentration and the fraction of white households without school-age children is much smaller than for white households with children. This is also generally consistent with the results reported in table 3, which showed that white households with no school-age children were unresponsive to the concentration of nonwhite students. Thus, the results with the aggregate data provide further support for our contention that among white households with children in schools with more nonwhite students, support for the voucher largely reflects a desire to use the voucher.

The results for elementary schools, reported in columns 3 and 4 of table 6, are also quite similar to those obtained using the survey data. For example, the point estimate on nonwhite concentration for white households with school-age children in column 3 is 0.177, compared to the survey data estimate of 0.139. Similarly, the point estimate on nonwhite concentration for nonwhite households with children in column 3 is −0.042 and statistically insignificant; the corresponding estimate from the survey data is −0.027, and is also statistically insignificant.34

As an additional robustness check of the survey results, we also estimated several expanded versions of equation (13). Specifically, we replicated the analysis in table 5 by separating nonwhite share into Hispanic, Asian, and black, and into LEP and non-LEP groups. Following the logic behind equation (13), these variables all enter our model as interaction terms. Results based on the expanded set of explanatory variables are reported in table 7.35 Analogous to table 5, the top panel of table 7 reports results for nonwhite share alone, panel 2 reports results for the three separate racial/ethnic groups, panel 3 shows LEP and non-LEP nonwhite shares, and panel 4 shows the complete breakdown for each race/ethnicity and LEP/non-LEP share.

In almost all cases, the results for white households with children are strikingly consistent with the results from the survey data. For example, in panel 3 of table 7, the coefficient on LEP nonwhite share is 0.474, and the coefficient on non-LEP nonwhite share is 0.119; using the survey data, the same variables have coefficients of 0.540 and 0.140. Similarly, the results reported in panel 5 of table 7 indicate that once we control for student performance on the API, the estimated coefficient on the nonwhite share declines sharply in magnitude and becomes statistically insignificant, just as we saw with the survey data. The two results that look different in the aggregate data are the coefficients on the share of black students and the share of Asian students. In the survey data, the coefficients on black share in panels 2 and 4 of table 5 are positive but never statistically significant. In table 7, these coefficients are once again positive but much larger in magnitude and statistically significant.

34 The coefficients on the other control variables reported in table 6 are also generally consistent with those obtained using the survey data. For example, similar to the results reported in table 2, the coefficients on fraction home owner and fraction female are negative, while the coefficient on the fraction of registered Republicans is positive.

35 Since the inclusion of these additional variables had little impact on the other coefficients in our model, table 7 reports only the coefficients on the interaction terms between the school-level race/ethnicity and LEP variables and the fraction of white households with children and the fraction of nonwhite households with children.
Table 8.—Estimated Coefficients for Households with Children Under Age Six

<table>
<thead>
<tr>
<th>Regression Specification</th>
<th>(1) Closest High School</th>
<th>(2) Closest Elementary School</th>
<th>(3) Three Closest Elementary Schools</th>
</tr>
</thead>
<tbody>
<tr>
<td>Nonwhite Share × (White with School-Age Children)</td>
<td>0.194**</td>
<td>0.142**</td>
<td>0.154**</td>
</tr>
<tr>
<td></td>
<td>(0.045)</td>
<td>(0.025)</td>
<td>(0.026)</td>
</tr>
<tr>
<td>Nonwhite Share × (White with Children Under 6)</td>
<td>0.132</td>
<td>0.179**</td>
<td>0.210**</td>
</tr>
<tr>
<td></td>
<td>(0.108)</td>
<td>(0.069)</td>
<td>(0.074)</td>
</tr>
<tr>
<td>Nonwhite Share × (Nonwhite with School-Age Children)</td>
<td>−0.075</td>
<td>−0.011</td>
<td>−0.024</td>
</tr>
<tr>
<td></td>
<td>(0.047)</td>
<td>(0.028)</td>
<td>(0.029)</td>
</tr>
<tr>
<td>Nonwhite Share × (Nonwhite with Children under 6)</td>
<td>−0.262**</td>
<td>−0.303**</td>
<td>−0.249**</td>
</tr>
<tr>
<td></td>
<td>(0.145)</td>
<td>(0.089)</td>
<td>(0.092)</td>
</tr>
<tr>
<td>Nonwhite Share × (White No Children)</td>
<td>0.019</td>
<td>0.016*</td>
<td>0.020**</td>
</tr>
<tr>
<td></td>
<td>(0.014)</td>
<td>(0.008)</td>
<td>(0.008)</td>
</tr>
<tr>
<td>Nonwhite Share × (Nonwhite No Children)</td>
<td>0.042</td>
<td>0.040</td>
<td>0.048*</td>
</tr>
<tr>
<td></td>
<td>(0.051)</td>
<td>(0.026)</td>
<td>(0.028)</td>
</tr>
<tr>
<td>Observations</td>
<td>25,715</td>
<td>38,066</td>
<td>38,066</td>
</tr>
<tr>
<td>R²</td>
<td>0.69</td>
<td>0.66</td>
<td>0.67</td>
</tr>
</tbody>
</table>

Robust, clustered standard errors in parentheses. All specifications include regional fixed effects and control variables listed in table 6. *Significant at 10%. **Significant at 5%.

C. Households with Children under Age Six

In all of the results presented thus far, we have focused on the voting behavior of households with school-age children and ignored the voting behavior of households with children who are not yet of school age. When our analysis was based on survey data, this omission was out of necessity since the PPIC survey asked respondents only whether there were school-age children in the household. This is an unfortunate limitation since we might expect the voting behavior of households with very young children to look less like the voting behavior of households without children and more like that of households with school-age children. Fortunately, our aggregate data allow us to examine that possibility. The 2000 census contains block-level data on the number of households with children 6 years of age or younger. We used those data to calculate the fraction of white and nonwhite households with children under age 6 in each block group or pseudo-block group. We then estimated an expanded version of equation (13) that included these additional two variables and interactions of these two variables with nonwhite concentration.

Results based on our expanded specification are reported in table 8. The results strongly suggest that white households with children under age 6 voted very much like white households with school-age children. In particular, in columns 2 and 3, where nonwhite concentration is measured at the elementary level, the estimated coefficients on the interaction between nonwhite concentration and the fraction of white households with children under age 6 are positive and statistically significant. In fact, at the elementary level, those results are stronger for white households with children under age 6 than for white households with school-age children. That finding is consistent with the notion that families with very young children are likely to be more concerned with the racial and ethnic composition of elementary schools than other families with school-age children. That conclusion is reinforced by the results reported in column 1, which suggest that relative to white households with school-age children, white households with children under age 6 are less concerned about the nonwhite concentration in their assigned high school.

V. Conclusion

The debate over school vouchers has many dimensions and will no doubt continue into the foreseeable future. In this paper, we provide evidence on one consequence of universal vouchers: their impact on racial and ethnic segregation across schools. In our empirical work, we find that among white households with children, support for vouchers increases with the proportion of nonwhite students in the local public schools. This is not true of nonwhite households, suggesting that if a voucher program were implemented, schools that currently have higher shares of nonwhite students would likely lose more white students,
and stratification would increase. The magnitude of the estimated effect appears to be nontrivial: in schools with a nonwhite share of 71% (the average for nonwhite households), voucher support among white voters is roughly 6 percentage points higher than in schools with a nonwhite share of 45% (the average for white households). This effect would likely be most pronounced in schools with higher proportions of students who are Hispanic.

It is important to point out that these conclusions are based on evidence from a single state, namely California. This is both a strength and weakness of the study. On the one hand, California’s size and diversity provides an ideal setting to examine how variation in the racial/ethnic composition of schools affects support for school vouchers. On the other hand, California has some unique characteristics that may make extrapolation of our results to other states difficult. For example, California has experienced a large increase in nonwhite immigration over the past several decades and that immigration has dramatically altered the ethnic composition of public schools. This may have had a profound effect on the voting behavior of white households with school-age children. In addition, California has one of the most centralized systems of public school finance in the nation. Due to court-ordered reforms and Proposition 13, California has transferred the responsibility for funding its schools from local school districts to the state and has equalized spending per pupil across districts. As a result, local school districts have almost no control over local school spending. This lack of local control may also have influenced voting behavior on Proposition 38.

It is also important to point out that the design of a voucher policy may affect the resulting behavior of households. The voucher policy proposed in Proposition 38 was a universal voucher, available to all students, while a voucher program that is targeted to a specific population may have a much smaller impact on stratification. At the same time, our results raise concerns about programs targeted to low-performing schools, as the targeted schools are precisely the ones that white households are most likely to want to leave. The value of the voucher also plays a role. The voucher amount under Proposition 38 was relatively low, and at most private schools, households would have needed to supplement the voucher in order to use it. As Ferreyra (2007) shows, voucher use among low-income households increases with the value of the voucher so with larger vouchers, differences between white and nonwhite households may be reduced.

Finally, our results give some indication that it is not directly race or ethnicity that is driving voter behavior but other school characteristics, such as limited English proficiency or student performance. Specifically, in the survey data, our results are driven by the voting preferences of whites in schools with substantial Hispanic populations, and when we further divide nonwhite shares by limited English proficiency, we find that our results are concentrated in schools with high shares of LEP Hispanic students. As Betts and Fairlie (2003) pointed out, white flight from limited English proficient students could be at least partially explained by resource constraints: teaching limited English proficient students requires additional resources (specially trained teachers, different curricula, and so on), and to the extent that schools are not fully compensated for these additional costs, this will divert resources from white students. This would suggest that one way for policymakers to mitigate the potential for segregation under vouchers would be to increase resources to schools with large concentrations of non-white LEP students.

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