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Supply vs. demand under an affirmative action ban: Estimates from UC law schools[☆]

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ABSTRACT

Affirmative action bans can reduce black enrollment not only by reducing black admission advantages (contracting demand) but also by reducing applications (contracting supply) from black students who can still gain admission but prefer alternative schools that still practice affirmative action. When affirmative action was banned at UC law schools, Berkeley's black applications and enrollment declined by almost half even as black admission rates *rose* relative to whites. I ask whether black enrollment at UC law schools would have markedly declined even if black supply had not contracted. I find in a large sample of students applying to law schools nationwide that black supply contractions were driven mostly or entirely by students unlikely to gain admission under the ban, yielding stronger post-ban black applicant pools. Holding applicant pools constant, I estimate that the ban reduced black admission rates at both Berkeley and UCLA by half. Hence, black enrollment would likely have plummeted even if black supply had not contracted—as could occur under a nationwide ban that eliminates affirmative-action-practicing alternatives.

1. Introduction

Black students in the United States would be substantially underrepresented at elite universities if admission decisions were made purely on the basis of academic credentials and without regard to race (Bowen and Bok,2000; Kane,1998; Espenshade et al., 2004; Rothstein and Yoon, 2008). Elite universities that value diversity therefore practice affirmative action: awarding admission advantages to black applicants on the basis of race. However, affirmative action is in legal jeopardy: seven states have banned the practice at their public universities, and the U.S. Supreme Court has indicated that it expects to broaden these bans nationwide by 2030.[1]

Affirmative action bans can reduce black enrollment through two related channels. First, affirmative action bans increase the opportunity cost of admitting black students by weakening the racial information that schools can use in admissions and thereby increasing the non-racial student strength that schools must forgo. The higher opportunity cost can induce schools to contract *demand* in the form of reduced black admission advantages (Chan and Eyster, 2003; Fryer et al., 2007; Epple et al., 2008).[2] Second, a demand contraction of any size can reduce the value to black students of attending an affected school (e.g. due to smaller black campus communities), inducing a *supply* contraction in the form of reduced applications even from black students who can still gain admission but prefer alternative

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[1] The most recent Supreme Court decision on affirmative action (*Grutter v. Bollinger* 2003) concluded with the widely quoted warning "We expect that 25 years from now [in 2028], the use of racial preferences will no longer be necessary to further the interest approved today" because "race-conscious admissions policies must be limited in time." The Court has just heard new oral arguments in the affirmative action challenge *Fisher v. Texas* (December 2015).

[2] See Appendix A.i for a simple model.

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schools that still practice affirmative action (Long, 2004; Card and Krueger, 2004, 2005; Dickson,2006).[3] Separating supply and demand effects can be crucial for modeling consequences of a nationwide affirmative action ban that eliminates affirmative-action-practicing alternatives and thus can mute supply responses (Arcidiacono, 2005; Epple et al., 2008). I separate these effects in the context of the first and largest ban—the 1996 University of California affirmative action ban—at the UC's elite law schools Berkeley and UCLA, which experienced extraordinary contractions in black applications after the ban.[4]

Fig. 1 motivates the analysis using public aggregates. It shows that after a transition period, the ban permanently reduced the black share of Berkeley's applicant pool by 47.7% even as black admission rates *rose* slightly relative to white admission rates—resulting in Berkeley's black enrollment share falling by 40.0% (as yield rates changed little).[5] The demand-centric explanation of these effects would be that Berkeley substantially contracted demand for black students and that only less-credentialed black students stopped applying, due to expected rejection. A nationwide ban would therefore also be expected to substantially reduce elite black enrollments. In contrast, the supply-centric explanation would be that Berkeley barely contracted demand for black students, but black students of all credential levels nevertheless stopped applying in favor of "black-friendlier" schools with slightly higher racial diversity. A nationwide ban may therefore have little effect on elite black enrollments, as no black-friendlier schools would exist. This paper asks: would UC black enrollments have markedly declined even if black students had not stopped applying?

I address this question using a large sample of applicants to UC and non-UC law schools. I find that black supply contractions were very concentrated among students unlikely to gain admission under the ban, yielding stronger post-ban black applicant pools. After controlling for selective attrition from applicant pools, I robustly estimate that the ban reduced the black admission rate in this sample by half at both Berkeley and UCLA. Hence based on this sample, black enrollment at these elite schools would likely have declined dramatically even if black students had not stopped applying.

Economically, one can understand the results as follows. Affirmative action bans weaken the racial information that can be used in admissions, which increases the non-racial student strength that schools must forgo in order to admit each additional black student. UC schools responded to this higher opportunity cost by collecting race-correlating information like diversity essays and maintained a selection-corrected black admission rate (31%) well above the rate that would prevail under pre-ban white admission standards (8%). But UC schools nevertheless substantially contracted demand for black students: the 31% selection-corrected post-ban black admission rate was still only half the 61% pre-ban rate. On the supply side, highly credentialed black students continued to apply (exhibiting no significant change) relative to less-credentialed black students (exhibiting a −42% change). This pattern is consistent with supply responding less to campus racial diversity than to one's own admission probability, with fixed per-school application costs. Colloquially, the results are consistent with black students still wanting to attend UC schools but simply not being able to get in anymore.

The findings are local to and made possible by administrative application-level data on all 25,499 applications submitted to law schools nationwide between 1990 and 2006 by 5,353 undergraduates from one elite college. The dataset's information on the application behavior of non-UC applicants *and* on the admission decisions of both UC and non-UC applicants permits this paper's joint analysis of supply and demand under the UC ban. The data contain only 185 applications of black students to UC schools, but the key specifications are nevertheless sufficiently statistically powerful because law school admissions are unusually formulaic and because effects are large. Omitted variables bias is possible but minimized in this context because the observed covariates are such powerful predictors of admission and because the dataset's thousands of independent screens—admission decisions of UC applicants at *non-UC* schools—provide a unique opportunity to control for an inferred measure of not-directly-observed applicant strength (e.g. recommendation letters), similar to Dale and Krueger (2002).

The results contribute to a large empirical literature on affirmative action bans. On the supply side, Card and Krueger (2005) and Dickson (2006) find no enduring response of minority applications—neither overall nor among highly credentialed minorities—to California's and Texas's bans, respectively, based on high-school student SAT submission data.[6] My paper studies a context with a huge and enduring overall black application response, which therefore provides an especially ripe opportunity to identify supply responses that could be driving large black enrollment declines. On the demand side, Long and Tienda (2008), Arcidiacono et al. (2014), and Antonovics and Backes (2014) also use administrative data to study admissions under an affirmative action ban. I break from their work by using data on students who applied to both affected and unaffected schools in order to study both supply and demand as well

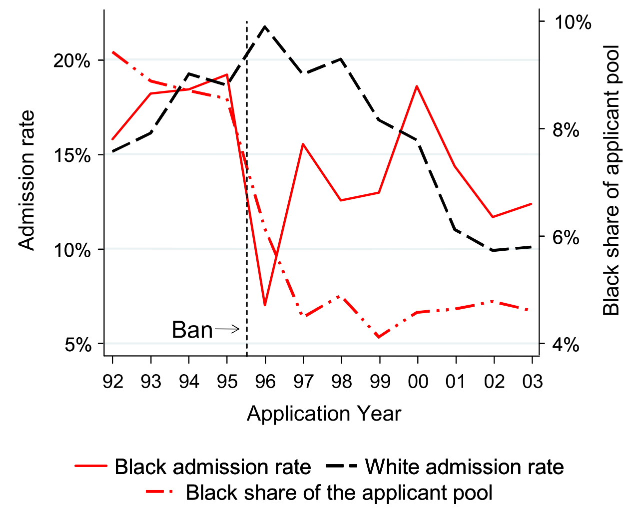

Fig. 1. Berkeley admission rates and racial mix of applicants. Notes — This graph uses public aggregates on the universe of applicants to Berkeley to plot the time series of overall admission rates by race and the black share of the applicant pool at Berkeley for all available years. Application year refers to the autumn of the application year. Averaged across 1992–1995 and 2000–2003, the black-minus-white admission rate differential rose from 0.6 to 2.6 percentage points, and the black share of the applicant pool declined from 8.9% to 4.7%. Yield rates (the share of admitted students who enrolled, not plotted here) changed comparably for blacks and whites: 29.2% to 32.9% for blacks and 30.4% to 33.1% for whites. The share of first-year enrollees who were black (not plotted here but implied by these preceding statistics) declined from 8.7% to 5.2%. These school–year–race aggregates contain no information on applicant credentials. See Online Appendix Fig. 1 for the analogous graph for UCLA. The underlying data are no longer posted on the website of the UC Office of the President but are available on the author's website.

3 See Appendix A.ii for the simple model of Card and Krueger (2004).

4 All references to Berkeley and UCLA pertain to their law schools.

5 Changes are measured from 1992–1995 to 2000–2003. Yield rates (the shares of admitted students who enrolled) did not change differentially across races (see the figure notes). UCLA exhibited similar declines in the black share of enrollees (48.7%) and the applicant pool (38.3%) though black admission rates fell relative to white admission rates, underscoring the limited informativeness of public aggregates; see Online Appendix Fig. 1.

6 Long (2004) finds no overall response to Texas' ban but a negative overall response to California's ban in 1999, which Card and Krueger find was small (−1.3%) and short-lived (insignificant and near-zero in 2000 and 2001).

as to address selection on unobserved characteristics like recommendation letter strength. I also study professional school admissions, which spawned the two landmark Supreme Court cases upholding affirmative action (*University of California v. Bakke* 1978; *Grutter v. Bollinger* 2003).

Finally, the selective response of black applications provides a key empirical moment to match in structural simulations of a nationwide affirmative action ban at law schools and related settings (Arcidiacono, 2005; Epple et al., 2008). My admission results indicate that such simulations should assume neither full elimination of black admission advantages—as assumed in Krueger et al. (2006) and Rothstein and Yoon (2008)—nor fully sustained advantages—as assumed in Fryer et al. (2007). A midpoint between those extremes may be more reasonable.

The remainder of the paper is organized as follows. Section 2 describes the UC affirmative action ban. Section 3 introduces the data. Section 4 presents the results. Section 5 concludes.

2. Legal and institutional environment

2.1. Legal environment

On November 5, 1996, California became the first state to ban affirmative action—awarding admission preference to underrepresented minorities on the basis of race—when voters approved Proposition 209 to amend the state constitution to read: "The state shall not discriminate against, or grant preferential treatment to, any individual or group on the basis of race, sex, color, ethnicity, or national origin in the operation of public employment, public education, or public contracting." In particular, no University of California applicant is to be preferred to another on the basis of race. The ban went into effect immediately at UC law schools.[7] Six other states have since passed similar bans.[8]

Legally, the UC affirmative action ban prohibits the use of race in choosing among applicants but permits the use of applicant characteristics that correlate with race as long as those characteristics have defensible non-racial justification if challenged in court. For example, UC schools are free to use of low family income (which correlates with black status) because broadening socioeconomic access is considered to be independently valuable to universities, but the use of participation in a black-focused extracurricular group would almost certainly be considered illegal. Law school admission decisions are made by a small group of selectors applying subjective criteria with little transparency, so the actual information used is unknown. UC schools (which refer throughout this paper to Berkeley and UCLA law schools) were not bound by any other new laws.[9]

Nationally, the U.S. Supreme Court in 5–4 rulings in both 1978 (*Regents of the University of California v. Bakke*) and 2003 (*Grutter v. Bollinger*) upheld the federal constitutionality of affirmative action, keeping the practice legal at all public universities not subject to a statewide ban. The Court's rationale is that although the U.S. Constitution guarantees equal protection to all races under the law, "the educational benefits that flow from a diverse student body" are a "compelling governmental interest" that justifies the use of race when there are no "workable race-neutral alternatives that will achieve the diversity the university seeks" (*Grutter*). However, the Court concluded *Grutter* with the widely quoted warning "We expect that 25 years from now [in 2028], the use of racial preferences will no longer be necessary to further the interest approved today" because "race-conscious admissions policies must be limited in time." The Court recently heard oral arguments in *Fisher v. Texas*, widely reported to bode poorly for affirmative action's future.[10]

Affirmative action is currently legal at all private universities but affirmative action may in principle be banned there too, such as through restrictions on all federal-aid-receiving universities. Perhaps as a result, most of the nation's top private universities petitioned the Court in 2003 in detailed amicus briefs to keep it legal at public universities.

2.2. Institutional responses

A large theoretical literature predicts that affected schools may respond to an affirmative action ban by shifting admission weight to legal black-correlates, at least partially sustaining pre-ban black enrollment levels (see Appendix A). Consistent with that prediction, UC application forms changed immediately after the ban. Beginning in 1996, application forms have stated that race is not a criterion for admission, and the page requesting applicant race has been diverted to a UC statistical department and not reported to admission offices. Application forms instead solicited new written information that correlates with race (law school applicants are rarely interviewed). For example before the ban, Berkeley gave applicants ten short unconnected prompt options for the personal statement, eight of which did not refer to diversity or disadvantages. Immediately after the ban and ever since, all ten were replaced by a single lengthy one that invited applicants to discuss their contributions to "the diversity of the entering class" and their backgrounds including "a personal or family history of cultural, educational, or socioeconomic disadvantage" (see Online Appendix Fig. 2). In 1998, Berkeley added a full-page socioeconomic questionnaire to its application form requesting information such as college attendance rates of high-school friends and whether the applicant was raised by a single parent. Beginning in 2001, UCLA solicited declarations of interest in a Critical Race Studies program and instituted admission preference for interested applicants.

The schools' diversity preferences likely changed little after the ban. UC administrators strongly opposed the ban before it passed and were not systematically replaced after it passed. As the California political climate turned against affirmative action in 1995, the UC president, UC vice-presidents, and the chancellor of each UC campus united to "unanimously urge, in the strongest possible terms," the continuation of affirmative action.[11] Berkeley's dean added "The need to diversify the legal profession is not a vague liberal ideal: it is an essential component to the administration of justice."[12] The day after voters approved the ban, the UC president announced that the question facing the university was "How do we establish new paths to diversity consistent with the law?"[13] One year after the ban, Berkeley's dean launched an audit of policies and procedures "to see

[7] The ban went into effect one year later at UC undergraduate campuses. The state constitutional amendment superceded the 1995 UC Board of Regents SP-1 resolution, which would have ended affirmative action beginning in 1997 and was later repealed to no legal effect.

[8] The six other states currently under affirmative action bans are Arizona, Florida, Michigan, Nebraska, New Hampshire, and Washington. Georgia and Texas had temporary bans.

[9] Soon after the ban, the State of California guaranteed that high school seniors graduating in the top 4% of their high schools would gain admission to at least one UC campus but not necessarily the one of their choice. With eight UC campuses, this had little binding effect on undergraduate admissions at the elite campuses of Berkeley and UCLA. No such guarantee applied to law school admissions.

[10] See for example the article "Supreme Court Justices' Comments Don't Bode Well for Affirmative Action" (http://www.nytimes.com/2015/12/10/us/politics/supreme-court-to-revisit-case-that-may-alter-affirmative-action.html).

[11] 1995 "Statement Supporting Affirmative Action by UC President, Chancellors, and Vice Presidents", http://www.development.umd.edu/Diversity/Response/Action/policy.

[12] 1995 press release, http://www.berkeley.edu/news/berkeleyan/1995/0524/regents.html.

[13] 1996 "Letter from President Richard C. Atkinson to the University Community Re: Passage of Proposition 209", http://www.universityofcalifornia.edu/news/article/20607.

whether [Berkeley] can achieve greater diversity" after "dire" admission results.[14] Berkeley's dean and the UC president continued in their posts through 2000 and 2003, respectively. Christopher Edley, a vocal proponent of affirmative action and adviser to President Bill Clinton on the topic, served as Berkeley's dean from 2004 to 2013. Other institutional features like the number of first-year enrollees remained nearly unchanged.

3. Data

3.1. Source, basic variables, and sample restrictions

This paper's primary dataset—which I call the Elite Applications to Law School (EALS)—comprises administrative application-level data on 67% of an elite college's seniors and graduates (collectively referred to here as "students") who applied to law schools nationwide between the fall of 1990 and the fall of 2006. Applications to every U.S. law school are submitted through the Law School Admissions Council, which records application information and the admission decision for every application filed.[15] Two-thirds of applicants choose to release their data to their colleges' administrators, and I obtained and digitized seventeen years of a single college's data. The college is elite, is not on the west coast, and has never been subject to an affirmative action ban. Subsection 3.2 investigates possible selection over time into the EALS, and Section 4 estimates and accounts for selection over time into the Berkeley and UCLA applicant pools.

The EALS contains six variables for each application: student race, LSAT test score (integers between 120 and 180), undergraduate grade point average (GPA) to two decimal places on a 4.00 scale, application year, law school submitted to, and admission decision. I standardize LSAT and GPA to each have mean zero and standard deviation one across students. Motivated semi-parametrically in Subsection 3.3 and used below in Fig. 3, I summarize applicants' LSAT and GPA scores with a scalar measure that I call "academic strength" equal to the standardized sum of standardized LSAT and standardized GPA, similar to the rescaling that Kling et al. (2007) employ in a different context. Application years 1990–1991 through 2001–2002 as well as 2005–2006 also contain applicant state of permanent residence; for these years, I digitized a California resident indicator for Berkeley and UCLA applications only.

The raw data contain 38,200 applications of 6,072 applicants to 187 law schools. For simplicity I restrict the analysis sample to the 94.3% of applicants listed as white, Asian, black, or Hispanic and the 78.9% of applications submitted to UC Berkeley, UCLA, or one of the fifteen most-applied-to schools that were never subject to an affirmative action ban. These fifteen schools correspond closely to the top-ranked law schools according to *U.S. News and World Report*, so I refer to them only somewhat imprecisely as the "top fifteen non-UC law schools."[16] The 170 other schools received relatively few applications in the EALS and are poor control schools for Berkeley and UCLA because these 170 other schools are less selective. The final seventeen-school EALS sample comprises 25,499 applications submitted by 5,353 applicants. Results reported in the main text restrict to the 17,814 applications from only the 3,774 black or white applicants; the appendix reports results using all races. See Appendix B for additional data-coding details.

3.2. Summary statistics

Table 1 lists summary statistics. The EALS sample is 61% white, 10% black, 19% Asian, and 10% Hispanic. Black applicants on average possess LSAT scores and GPA's 1.1 and 1.0 standard deviations lower, respectively, than white applicants. Online Appendix Fig. 3a-c use non-parametric densities of these academic characteristics to illustrate the first-order stochastic dominance of the black and Hispanic distributions by the white and Asian distributions. This stochastic dominance motivates universities' use of affirmative action in order to achieve more racially diverse cohorts. Online Appendix Fig. 3d plots means of academic strength over time by race among EALS applicants; pre-ban and post-ban means are very similar within races, suggesting little differential selection over time into the EALS. Section 4 estimates and accounts for differential selection over time into the Berkeley and UCLA applicant pools.

Berkeley received applications from 28% of all applicants (1,594 making it the seventh-most-applied-to school in this sample) and UCLA received applications from 14% of all applicants (777, the thirteenth most in this sample); see Online Appendix Table 1 for additional comparisons. These schools received relatively few applications from black students—60 before the ban and 67 after the ban at Berkeley, and 31 before the ban and 27 after the ban at UCLA—which is unsurprising given the relatively small size of elite professional school cohorts. The EALS nevertheless provides sufficient statistical power because law school admission decisions are largely determined by academic credentials and race and because effects are large.

3.3. Race and admission in the pre-ban cross section

To provide a feel for the admission data and also to motivate the use of a standardized measure of academic strength below in Fig. 3, Fig. 2a displays the semi-parametric relationship between LSAT, GPA, and admission within race–school–years in the EALS, using a 5% random sample of all 23,128 applications submitted to non-UC schools (Online Appendix Fig. 4 displays the 100% sample, intelligible only in color). Each application's admission decision is plotted in (LSAT, GPA) space, where each application's LSAT score has been re-centered by the estimated race–school–year fixed effect in order to account for selectivity differences across races, schools, and years. Specifically I fit a probit regression of admission on standardized LSAT (mean zero and standard deviation one), standardized GPA, and school–year–race fixed effects; add each application's estimated school–year–race effect to its LSAT value; and plot individual application decisions in GPA vs. adjusted LSAT space. Applications above and to the right of the best-fit admission threshold line have high enough LSAT and GPA scores to have a predicted admission probability of more than 50%, while those below and to the left do not.[17]

The best-fit line correctly predicts 89.1% of all admission decisions, and incorrect predictions are concentrated near the line. The ratio of the coefficients on LSAT and GPA in the underlying probit is 0.95, indicating that a one standard deviation higher LSAT is about as valuable in the admissions cross section as a one standard deviation higher GPA. When useful for subsequent illustrations, I therefore summarize an applicant's academic strength as the standardized

[14] 1997 Berkeley press release, http://berkeley.edu/news/berkeleyan/1997/0820/kay.html.

[15] Academic credentials are verified through third-party reports, and race is reported by applicants. Dishonest answers are grounds for revocation of an admission offer, expulsion from law school, or disbarment. To the extent that any applicants misreported their race, the EALS race variable nevertheless represents the race that was reported to schools on application forms.

[16] Deviations from *U.S. News* rankings are usually explained by a lower-ranked school being located in a large city. Berkeley was ranked sixth and UCLA was ranked fifteenth in 2010.

[17] The probit model is $Pr(ADMITTED_{istr}) = \Phi(\beta_1 LSAT_i + \beta_2 GPA_i + \gamma_{str})$ where i denotes an applicant and γ_{str} denotes the school–year–race fixed effects. Adjusted LSAT equals $LSAT_i + \hat{\gamma}_{str}/\hat{\beta}_1$. The slope of the best-fit admission threshold line is 0.95, equal to $-\hat{\beta}_1/\hat{\beta}_2$.

Table 1
Applicant Characteristics by Race.

	Share of applicants	LSAT score (sd 6.7)	Undergraduate GPA (sd 0.33)	Academic strength (mean 0, sd 1)	Admission rate
A. All applicants ($N = 5353$, collectively submitting 25,499 applications to top-17 schools)					
White	60.8%	167.3	3.47	0.24	41%
Black	9.7%	159.9	3.15	−0.98	56%
Asian	19.4%	167.6	3.52	0.33	41%
Hispanic	10.1%	162.8	3.31	−0.48	39%
B. Applicants to Berkeley ($N = 1594$)					
White	56.6%	167.5	3.47	0.23	31%
Black	8.0%	160.8	3.13	−0.92	43%
Asian	24.2%	167.0	3.49	0.21	36%
Hispanic	11.3%	162.3	3.31	−0.53	34%
C. Applicants to UCLA ($N = 777$)					
White	55.0%	165.4	3.38	−0.09	54%
Black	7.5%	159.6	3.03	−1.17	53%
Asian	24.5%	165.2	3.43	−0.06	60%
Hispanic	13.1%	159.8	3.23	−0.89	35%

Notes — Panel A lists mean student characteristics for the Elite Applications to Law School (EALS) sample used in this paper. The sample comprises the 5353 students who together submitted 25,499 applications over seventeen years to Berkeley, UCLA, and the top-fifteen law schools that were never subject to an affirmative action ban. LSAT is the standardized test score used in law school admissions and ranges from 120 to 180. Undergraduate grade point average is the cumulative undergraduate GPA on a 4.00 scale. Academic strength is a scalar index of the strength of an applicant's academic credentials, equal to the standardized (mean zero and standard deviation one) sum of standardized LSAT and standardized GPA (see Fig. 2) and is used only for Fig. 3. Panels B and C list the same statistics for applicants to Berkeley and UCLA, respectively, in the EALS. Online Appendix Table 1 lists summary statistics on application behavior and comparisons to the nationwide population of law school applicants.

(mean zero, standard deviation one) unweighted sum of standardized LSAT and standardized GPA. Fig. 2b confirms that the semi-parametric relationship between academic strength and admission within race–school–years is well-approximated by a univariate probit regression of admission on academic strength alone. I refer to such a curve relating admission to academic strength as an admission rule in academic strength.

Fig. 2c plots fitted admission rules for blacks and whites in pre-ban Berkeley and UCLA admissions.[18] For ease of comparison, each school's fitted rules have been shifted horizontally by an additive constant so that the admission probability for whites equals 0.5 at academic strength 0. The graph shows that there are levels of academic strength at each school where blacks were nearly assured admission and whites were nearly assured rejection. Berkeley's black and white admission rules are separated by 1.90 standard deviations of academic strength, implying black status is observed to be worth more than the difference between an A- GPA and a B- GPA for a given LSAT in the pre-ban cross section.[19] At UCLA, the difference is 1.39 standard deviations. Had pre-ban black applicants to each school been subjected to the observed pre-ban white admission standards, Berkeley's black admission rate is predicted to have been 6% rather than the actual 57%, and UCLA's to have been 10% rather than 65% (documented in Section 4.2 below). These black–white differences in the EALS are similar in magnitude to those found in the universe of law school applicants to elite schools like Berkeley and UCLA (Rothstein and Yoon, 2008) and in undergraduate admissions (Bowen and Bok,2000; Kane,1998; Espenshade et al.,2004).[20]

[18] For each school I estimate the probit model $Pr(ADMITTED_{it}) = \Phi(\beta_1 ACADEMICSTRENGTH_i + \beta_2 BLACK_i + \gamma_t)$ using pre-ban black and white applications, where $BLACK_i$ is a black indicator and γ_t denotes year fixed effects. This paper focuses on black outcomes for simplicity and statistical power, but results for Hispanics are similar.

[19] That is, $\hat{\beta}_2/\hat{\beta}_1 = 1.90$ in the underlying Berkeley regression.

[20] Using individual-level data on matriculants but not applications, Rothstein and Yoon estimate that black enrollment at elite law schools would have been 90% lower under white admission standards.

3.4. Inferred strength

The previous subsection showed that LSAT and GPA explain the vast majority of the variation in within-race admission decisions, and the difference-in-differences analysis below will hold applicant pools constant along LSAT, GPA, and race. However, one may yet be concerned in that analysis that there is selection on unobservables, conditional on LSAT and GPA across races and over time. In particular, all top law schools solicit and are believed to value additional applicant characteristics like recommendation letters, leadership experience, and a background of no criminal behavior or academic dishonesty.[21] I proxy for such commonly-valued unobserved admission determinants using the intuition that if an applicant who is predicted to be rejected based on LSAT, GPA, and race is in fact consistently admitted across schools in the EALS, then this applicant is likely strong on unobserved characteristics like recommendation letters.[22]

Specifically, I construct an inferred strength variable for each application, equal to the mean admission success that a given applicant experienced in her other applications that is not explained by observed characteristics. For each school s in either the pre-ban (1990–1995) or post-ban (1996–2006) era, I fit:

$$\Pr(ADMITTED_{ist}) = \Phi(\beta_1 LSAT_i + \beta_2 GPA_i + \beta_3 BLACK_i + \beta_4 HISPANIC_i + \beta_5 ASIAN_i + \gamma_t)$$

where $ADMITTED_{it}$ is an indicator for whether student i's application in year t earned an admission offer; $BLACK_i$, $HISPANIC_i$, and $ASIAN_i$ are indicators of racial status; γ_t is a vector of year fixed effects; and $\Phi(\cdot)$ denotes the Normal cumulative distribution function using only the applications submitted to school s in the given era. I use the resulting coefficients to compute a predicted admission probability

[21] Admission selection criteria are highly correlated across law schools; Fig. 2a showed this to be the case for directly observed applicant characteristics (LSAT, GPA, and race). Characteristics that are valued inconsistently across admission offices include the applicant's geographic preference and intended legal specialty.

[22] Dale and Krueger (2002) similarly use the rich information embedded in independent screens (admission decisions at other schools) to estimate the returns to higher education.

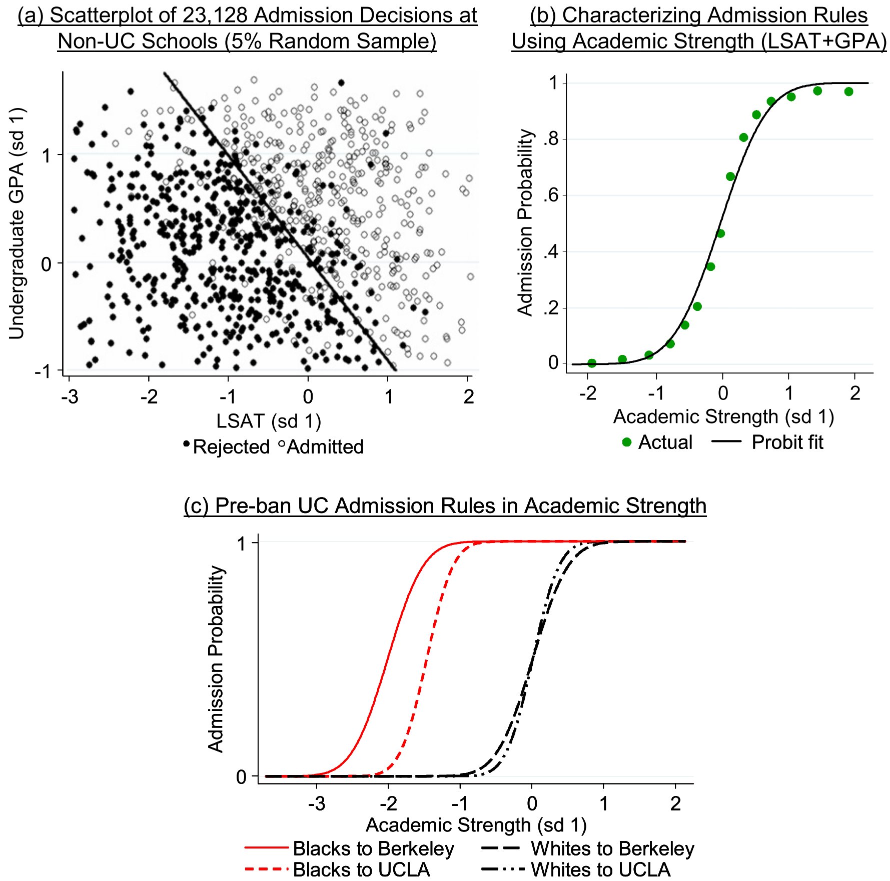

Fig. 2. Race, academic credentials, and admission under affirmative action. Notes — Panel A plots standardized LSAT score (mean zero and standard deviation one), standardized undergraduate GPA, and the actual admission decision for a 5% random sample of the 23,128 Elite Applications to Law School (EALS) applications submitted to the top-fifteen non-UC schools that were never subject to an affirmative action ban. Online Appendix Fig. 4 displays the full sample in color. To account for cross-school selectivity differences, each application's LSAT has been additively shifted by its school–year–race fixed effect from a probit regression of admission on LSAT, GPA, and these fixed effects (see Section 3.3); the overlaid best-fit admission threshold line correctly predicts 89.1% of admission decisions. The regression indicates that a one standard deviation higher LSAT is about as valuable in the admissions cross section as a one standard deviation higher GPA. Thus when useful for Fig. 3, I summarize an application's LSAT and GPA with the scalar index academic strength, equal to the standardized sum of standardized LSAT and standardized GPA. Panel B plots admission rates within fifteen academic strength bins using all 23,128 non-UC applications and overlays the univariate probit fit, where each application's academic strength has been additively shifted by its school–year–race fixed effect from a probit regression of admission on academic strength and these fixed effects. Panel C plots probit-fitted admission rules by race at UC schools before the 1996 affirmative action ban, derived from a regression of admission on academic strength, a black indicator, and year fixed effects using pre-ban black and white applications to Berkeley, and separately for UCLA. For ease of comparison, each school's pair of admission rules has been horizontally shifted by an additive constant so that the predicted admission probability for whites equals 0.5 at academic strength 0.

$Pr\,(ADMITTED_{ist})$ for each application and compute admission residuals $\varepsilon_{ist} = ADMITTED_{ist} - Pr\,(ADMITTED_{ist})$ for each application. Then for each application *ist*, I compute inferred strength equal to the leave-out mean of student *i*'s admission residuals *from her applications to schools other than s*:[23]

$$INFERREDSTRENGTH_{ist} = \overline{\varepsilon_{is't}}, s' \neq s.$$

Note that this leave-out-mean formula uses information only from independent screens (admission decisions at schools other than *s*) to assign the inferred strength value for the applicant's application to school *s*. When using inferred strength in student-level regressions of the decision to apply to UC schools in Section 4.1, I compute each student's inferred strength as the average of inferred strength across the student's applications.

Inferred strength ranges from −1 to 1 and is positive for applications submitted by students with relatively weak direct observables who were nevertheless accepted at other schools. Likewise, inferred strength is negative for applications submitted by students with relatively strong observables who were nevertheless rejected at other schools.[24] Online Appendix Fig. 5 demonstrates the predictive power of the inferred strength variable using the full sample of applications by plotting the strongly upward-sloping non-parametric relationship between admission and inferred strength, conditional on LSAT, GPA, and school–year fixed effects.

[23] That is, $INFERREDSTRENGTH_{ist} = \frac{1}{S_{it}-1}\sum_{s'=1,s'\neq s}^{S_{it}} \varepsilon_{is't}$, where S_{it} equals the total number of schools applied to by student *i* in year *t* and where the schools applied to by applicant *i* in year *t* are indexed 1 to S_{it}. To flexibly handle the small share of applicants who applied to only one school, I assign their applications inferred strength equal to zero and include an indicator for these applicants in all regressions where inferred strength is used.

[24] For example, consider a student who applied to Berkeley, Harvard, and Northwestern; who had an admission probability of 0.25 at Harvard and 0.75 at Northwestern based on her LSAT, GPA, and race and the selectivity at Harvard and Northwestern in the given application year; and who was admitted at both Harvard and Northwestern. This candidate's application to Berkeley would be assigned an inferred strength value of 0.5 (= [(1 − .25) + (1 − .75)]/2).

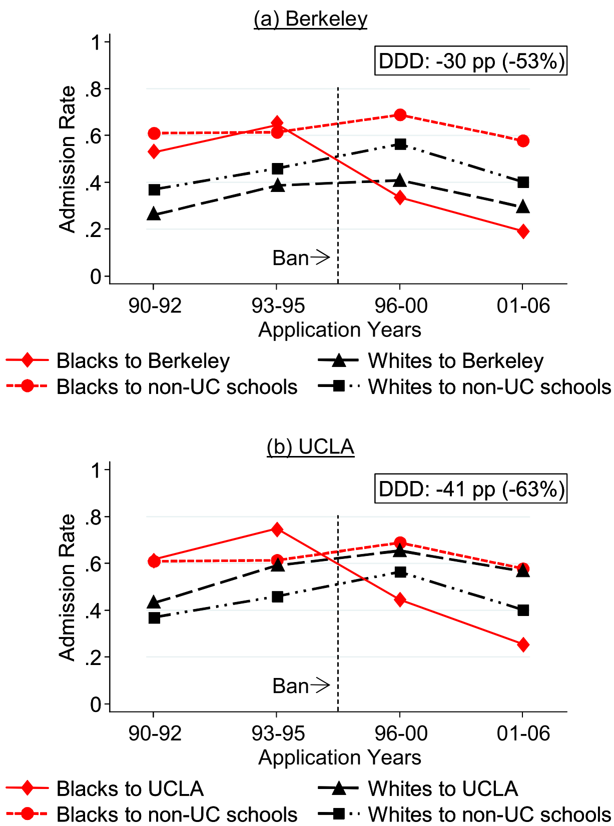

Fig. 3. Selection-corrected admission rates by race. Notes — This figure displays the time series of black and white admission rates at Berkeley, UCLA, and non-UC schools, where applicant characteristics have been held constant at pre-ban levels using semi-parametric reweighting as in DiNardo et al. (1996). To construct the time series of black admission rates at Berkeley, I first compute terciles of academic strength among pre-ban black applications to Berkeley. Then for each time period shown in the figure, I weight black applications to Berkeley so that each pre-ban-defined tercile receives equal weight when computing the displayed admission rate. I repeat this process for whites at Berkeley and for whites and blacks separately at UCLA and at each non-UC school, averaging across non-UC schools to construct the non-UC series. This semi-parametric reweighting on academic strength is data-demanding, so I group the data into two pre-ban time periods (1990–1992 and 1993–1995) and two post-ban time periods (1996–2000 and 2001–2006). Pooling all pre-ban years and all post-ban years, the triple-difference (DDD) estimate of the effect of the ban on the black admission rate at each UC school is overlaid, with the DDD estimate as a fraction of the pre-ban black admission rate in parentheses. Online Appendix Table 2 lists the numbers underlying these DDD estimates. Table 2 reports parametric DDD estimates that account for the minor space-opening effect of a decline in black admission rates.

4. Results

This section uses the EALS to estimate the effect of the UC affirmative action ban on applications to and admissions at Berkeley and UCLA. All estimates are local to the EALS. I begin by investigating the application (supply) response to the ban, finding that black student attrition from UC applicant pools was driven by less-credentialed black students. This implies that the UC's average black applicant has become more highly credentialed, so raw admission rate changes that do not control for selective attrition from applicant pools (like those shown in Fig. 1) can fail to reflect changes in black admission advantages. I then correct for selective attrition to estimate the paper's main object of interest: the admission (demand) response to the ban. I robustly find that the ban causes a large reduction in black admission advantages at UC schools. However, large observed cross-sectional black admission advantages remain.

4.1. Applications (Supply)

I test for effects of the UC affirmative action ban on the likelihood that EALS black students applied to each UC law school by fitting probit models based on the following DD specification:

$$\Pr(APPLIED_{it}) = \Phi(\mathbf{X}_i\alpha + \beta_1 BLACK_i + \beta_2 BLACK_i \times POST_t + \gamma_t) \quad (1)$$

using black and white students, where $APPLIED_{it}$ is an indicator for whether student i in year t submitted an application to the UC school being studied; $BLACK_i$ is an indicator for black racial status; $POST_t$ is an indicator for the application being submitted after the ban; $\mathbf{X}_i$ is a vector containing LSAT score, GPA, inferred strength, and potentially other covariates linearly, depending on the specification; γ_t is a vector of year fixed effects; and $\Phi(\cdot)$ denotes the Normal cumulative distribution function.[25] The coefficient β_2 is the coefficient of interest: the effect of the ban on the likelihood that a black student applied to the UC school being studied. Reported coefficients are marginal effects averaged over the right-hand-side characteristics of pre-ban black students, accompanied by robust standard errors (the dataset comprises one observation per student).

Table 2 column 2 presents the results for whether the applicant applied to Berkeley (panel A) or UCLA (panel B). Panel A reports that the ban reduced application rates to UC Berkeley among black students in the EALS by 9.3 percentage points with a t-statistic of 2.6 and equal to a 34.7% decline relative to the actual pre-ban mean among black students of 26.8 percentage points. Panel B shows an identical effect size in percentage terms (34.4%) at UCLA. These effect sizes are comparable to those shown for the full Berkeley and UCLA applicant pools (47.7% and 38.3%) in Fig. 1 and Online Appendix Fig. 1.

Columns 3–5 present results by whether black students could still be expected to be admitted with high probability—which cannot be studied using public aggregates. Column 3 replicates the regression underlying column 2 while including two additional covariates that divide students by a composite measure of applicant strength: an indicator for whether the student had at least a 99% predicted probability of admission to the given UC school (under pre-ban standards based on pre-ban estimation of Eq. (1) with admission as the dependent variable, as in column 7 introduced below), as well as the interaction between this "highly credentialed" indicator and the black-x-post-ban indicator.[26] Using pre-ban admission standards to categorize students has the property that the categorization is not endogenous to the policy change. For interpretation, highly credentialed post-ban black students (32.0% of all black EALS students for Berkeley and 43.5% for UCLA) were on average still quite likely to be admitted post-ban: a 67.3% admission probability at Berkeley and 91.8% at UCLA based on post-ban estimation of the column 7 regression. In contrast, non-highly-credentialed ("less-credentialed") post-ban black applicants had on average a 9.1% admission probability at Berkeley and 17.9% at UCLA.

The coefficients in panel A column 3 indicate that the large negative effect on black applications reported in column 2 was driven entirely by less-credentialed black applicants. Highly credentialed black applicants are estimated to have been insignificantly 2.7 percentage points more likely to apply post-ban. In contrast, less-credentialed black applicants were significantly 14.2 percentage points less likely to apply (t-statistic of 3.6), equal to a −48.0% change

[25] Results are similar when including Hispanics and Asians along with Hispanic and Asian indicators, or when omitting inferred strength. Application probabilities can be non-monotonic in the controls but including higher orders of the controls barely changes the results. Basic ordinary least squares are also reported.

[26] For this categorization, I assume that post-ban students were applying to law schools in 1992 (i.e. I use the 1992 fixed effect) which was an approximately average-selectivity pre-ban year.

Table 2
Effects on Black Application Rates and Selection-Corrected Black Admission Rates.

Dependent variable:	Applied					Admitted			
	(pp)	(pp)	(pp)	(pp)	(pp)	(pp)	(pp)	(pp)	(pp)
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
A. Berkeley									
Black × Post-ban	−10.5	−9.3	2.7	8.2	1.4	−39.7	−40.0	−33.9	−43.6
	(4.0)	(3.6)	(6.1)	(10.0)	(7.2)	(6.8)	(5.3)	(6.5)	(5.5)
Black × Post-ban × Pre-ban-admit-rate<99%			−14.2						
			(4.0)						
Black × Post-ban × Post-ban-admit-rate<90%				−15.5					
				(5.4)					
Black × Post-ban × Post-ban-admit-rate<75%					−11.5				
					(5.1)				
Ordinary least squares estimation	X					X			
Probit estimation		X	X	X	X		X	X	X
National trend controls								X	
CA residency control									X
Number of observations	3774	3774	3774	3774	3774	1029	1029	17,329	779
Number of clusters	3774	3774	3774	3774	3774	1029	1029	3754	779
Actual pre-ban black dependent variable mean	26.8	26.8	26.8	26.8	26.8	56.7	56.7	56.7	56.7
Δ Implied by Black × Post-ban effect						−34.2	−35.6	−30.0	−39.7
B. UCLA									
Black × Post-ban	−5.1	−4.8	−0.4	−0.9	−1.9	−48.1	−35.0	−33.5	−31.1
	(3.0)	(2.9)	(4.3)	(4.6)	(4.1)	(10.5)	(11.2)	(11.1)	(10.5)
Black × Post-ban × Pre-ban-admit-rate<99%			−6.1						
			(3.1)						
Black × Post-ban × Post-ban-admit-rate<90%				−5.0					
				(3.6)					
Black × Post-ban × Post-ban-admit-rate<75%					−4.4				
					(3.5)				
Ordinary least squares estimation	X					X			
Probit estimation		X	X	X	X		X	X	X
National trend controls								X	
CA residency control									X
Number of observations	3774	3774	3774	3774	3774	485	485	16,785	371
Number of clusters	3774	3774	3774	3774	3774	485	485	3736	371
Actual pre-ban black dependent variable mean	13.8	13.8	13.8	13.8	13.8	64.5	64.5	64.5	64.5
Δ Implied by Black × Post-ban effect						−41.6	−32.0	−30.2	−29.2

Notes — This table uses black and white students in the EALS to report difference-in-difference estimates of the effect of the UC affirmative action ban on whether a student applied to (columns 1–5) or was admitted at (columns 6–9) Berkeley (panel A) or UCLA (panel B). Standard errors clustered by student (equivalent to robust standard errors in all specification except column 8) are in parentheses, and reported coefficients in probit columns are marginal effects averaged over the UC school's pre-ban black applicant characteristics. All specifications include a black indicator, a black indicator interacted with a post-ban indicator, LSAT test score, undergraduate GPA, inferred strength, and year fixed effects; see Table 1 and Fig. 2 for variable definitions. Specifications with national trend controls are triple-difference regressions that control for black and white admission trends at non-UC schools; these regressions include school–year fixed effects and are weighted so that each school receives equal weight in each time period (pre-ban and post-ban). The California residency indicator control is available only for applications to UC schools and in certain years. Column 3 includes two covariates that divide applicants by a composite measure of applicant strength: an indicator for whether the applicant had a 99% predicted probability of admission to the given UC school based on pre-ban estimation of the column 7 regression, as well as the interaction between this indicator and the black-×-post-ban indicator. Columns 4–5 present alternative specifications that group black applicants into those who were quite likely (at least 90% for column 4 and at least 75% for column 5) to be admitted after the ban, based on post-ban estimation of the regression underlying column 7. The final row in each panel of columns 6–9 reports estimates of the change in the admission rate that pre-ban black applicants are predicted to have experienced had the ban been in effect, accounting for the minor space-opening effect of a decline in black admission rates (see Section 4.2 for details).

relative to this sub-group's pre-ban mean application rate of 29.6%. The standard error on the effect among highly credentialed black applicants does not permit rejection of all meaningful response magnitudes, but the heterogeneity in application behavior is clear. Panel B shows similar effects at UCLA, with a percentage change among less-credentialed black applicants of −35.0% (= −6.0/17.4). Thus, on average across Berkeley and UCLA, the ban reduced applications from less-credentialed black students by 41.5%. Columns 4–5 present alternative specifications that categorize black applicants into those who were quite likely (at least 90% for column 4 and at least 75% for column 5) to be admitted *after* the ban, based on *post-ban* estimation of the regression underlying column 2.[27] Results are qualitatively similar to those in column 3.

I conclude that there is robust evidence of a large decline in applications to UC schools from less-credentialed black applicants with no evidence of a decline in applications from highly credentialed black applicants. This implies that the average post-ban black applicant to UC schools was substantially more highly credentialed than the average pre-ban black applicant, relative to contemporaneous white applicants. Hence, raw changes in the black–white admission rate gap (like the one displayed in Fig. 1) can fail to reflect changes in black admission advantages (demand responses). The next subsection estimates the change in black admission advantages at UC schools by estimating the change in black admission rates, correcting for selective attrition from UC applicant pools.

4.2. Admissions (Demand)

For simplicity and transparency, I first display the time series of selection-corrected admission rates for black and white applicants at UC and non-UC schools using semi-parametric reweighting on academic strength. Fig. 3 displays the time series of black and white

[27] For these categorizations, I assume that pre-ban students were applying to law schools in 1999 (i.e. I use the 1999 fixed effect) which was an approximately average-selectivity post-ban year.

admission rates at Berkeley, UCLA, and non-UC schools, where applicant characteristics have been held constant at pre-ban levels using simple semi-parametric reweighting as in DiNardo et al. (1996). To construct the time series of black admission rates at Berkeley, I first compute terciles of academic strength among only pre-ban black applications to Berkeley. Then for each time period shown in the figure, I weight black applications to Berkeley so that each pre-ban-defined tercile receives equal weight when computing the displayed admission rate.[28] I repeat this process for whites at Berkeley and for whites and blacks separately at UCLA and at each non-UC school, averaging resulting admission rates across non-UC schools to construct the plotted non-UC series. This semi-parametric reweighting is data-demanding, so I reweight on academic strength only and group the data into two pre-ban time periods (1990–1992 and 1993–1995) and two post-ban time periods (1996–2000 and 2001–2006).

The figure shows that at non-UC schools, there was little change over time in the difference between black and white admission rates. At Berkeley the black admission rate rose between 1990–1992 and 1993–1995 about as much as the white admission rate did, thus exhibiting parallel pre-ban trends. Between 1993–1995 and 1996–2000, the black admission rate fell from 64.4% to 33.3% and did not subsequently recover relative to the white admission rate. Fig. 3b shows a similar decline at UCLA. One can use these reweighted admission rates to compute a simple selection-corrected triple-difference (DDD) estimate of the effect of the ban on the black admission rate at each UC school: −29.9 percentage points at Berkeley (relative to the actual pre-ban black admission rate of 56.7%) and −40.7 percentage points (relative to the actual pre-ban black admission rate of 64.5%).[29] These declines were much larger than those observed at any non-UC school, so the empirical p value on each of these declines relative to the distribution of changes at non-UC schools is 0.

Though parsimonious and transparent, Fig. 3 does not control for LSAT and GPA separately, does not control inferred strength or California residency, and does not allow for selection within tercile bins. Table 2 columns 5–9 report regression estimates of the effect of the ban on black admission outcomes at each UC school. The underlying regressions are based on Eq. (1) using black and white applications to a given UC school and use the dependent variable $ADMITTED_{it}$, which is an indicator for whether student i′s application in year t earned an admission offer. When producing DDD estimates that account for national trends, I include all black and white applications to the top-fifteen non-UC schools and interact the second and third terms with an indicator for the application being submitted to a non-UC school.[30] Standard errors are clustered at the student level. Online Appendix Tables 3–5 replicate Table 2 columns 6–9 using alternative specifications that include all races or control for more interactions.

Column 8 controls for national trends and is my preferred specification. Panel A reports that the ban caused an estimated 33.9 percentage point reduction in black applicants' probability of admission, averaged over the right-hand-side characteristics of pre-ban black applicants and relative to the actual pre-ban black admission rate of 56.7%. Panel B reports an analogous estimate for UCLA of −33.5 percentage points, relative to the actual pre-ban black admission rate of 64.5%. These estimates have t statistics of 5.2 and 3.0 respectively. The other columns report similar magnitudes in other specifications, including column 9 which controls for California residency—a UC-specific admission determinant that inferred strength is unlikely to encompass—in the years it is available.

A decline in black admission rates relative to whites opens up space in the admitted cohort for both black and white applicants, implying that the above estimates somewhat overstate the effect of the ban on black admission rates. I therefore compute an adjusted estimate of the effect of the ban on the black admission rate at each UC school by first using the UC-specific coefficients of each regression to compute a probit latent variable value for each black and white pre-ban application according to post-ban criteria. I then add a constant to every application's value until the mean predicted admission probability across applications equals the actual admission rate observed among these applications.[31] The resulting estimates are reported in the bottom row of each panel of Table 2 column 8: −30.0 percentage points at Berkeley and −30.2 percentage points at UCLA. These are my preferred estimates of the effect of the ban on UC black admission rates in the EALS. These declines were much larger than those estimated at any individual non-UC school, so the empirical p value of each of these estimates is 0. Averaging these DDD estimates across Berkeley and UCLA, I conclude that the ban reduced the black admission rate from 60.6% to 30.5% in the EALS when holding applicant pools constant.

Online Appendix Table 4 replicates Table 2 columns 6–9 using applications from all races (white, black, Hispanic, and Asian); the results are very similar to those in Table 2. Online Appendix Table 5 replicates Online Appendix Table 3 while also fully interacting covariates $\mathbf{X}_i$ with race indicators, the post-ban indicator, and the non-UC indicator; the DD results are somewhat larger in magnitude (more negative) than those in Table 2.

As a benchmark for the large effect sizes estimated above, I estimate the black admission rate that would prevail under observed pre-ban white admission standards—i.e. the black admission rate that would prevail if the ban simply eliminated cross-sectional black admission advantages. Specifically, I estimate the cross-sectional analogue to Eq. (1) for each UC school among pre-ban black and white applications:

$$\Pr(ADMITTED_{it}) = \Phi(\mathbf{X}_i\alpha + \beta_1 BLACK_i + \gamma_t) \tag{2}$$

where $\mathbf{X}_i$ is a vector of LSAT, GPA, and inferred strength and γ_t are year fixed effects.[32] I then use only the estimated coefficient vector $\hat{\alpha}$ and the year fixed effects to compute a probit latent variable value for each application. Finally to account for the fact that a decline in the black admission rate opens up space in the admitted cohort, I add a constant to every application's value until the mean predicted admission probability across applications equals the actual admission rate among these applications.

Columns 1–2 of Table 3A report the results. Whereas Berkeley actually admitted 56.7% of pre-ban black applicants, I estimate that it would have admitted only 5.6% under observed white admission standards. For UCLA, the statistics are 64.5% and 10.4%. Thus, averaging across Berkeley and UCLA, I estimate that the black admission

28 That is, each application in time period T with academic strength lying in tercile G receives weight $1/N_{TG}$, where N_{TG} is the number of applications in the sample submitted to Berkeley in time period T with academic strength in tercile G. Quartiles yield similar results; I use terciles because some bin counts are small.

29 Pooling pre-ban years and separately pooling post-ban years for each series, each DDD estimate is equal to the change in black admission rates at the UC school, minus the change in white admission rates at the UC school and the change in the black–white admission rate difference at non-UC schools. See Online Appendix Table 2 for the arithmetic.

30 The DDD specification is $Pr(ADMITTED_{ist}) = \Phi(\mathbf{X}_i\alpha + \beta_1 BLACK_i + \beta_2 BLACK_i \times POST_t + \beta_3 BLACK_i \times UC_s + \beta_4 BLACK_i \times POST_t \times UC_s + \gamma_{st})$, where UC_s is an indicator for whether the application was submitted to the UC school being analyzed and γ_{st} is a vector of school–year fixed effects. I weight applications so that each school carries equal weight in each time period (pre-ban and post-ban).

31 Adding a constant varies selectivity uniformly across applications (i.e. preserves application rank). I obtain similar results under the similar method of using the UC-specific coefficients to rank pre-ban applications and then admitting the N highest-ranked applications, where N equals the total number of black and white pre-ban EALS applicants that the UC school admitted.

32 Results are similar when omitting inferred strength or including Hispanics and Asians in the regression along with Hispanic and Asian indicators.

Table 3
Black-White Admission Rate Differences in Pre-ban and Post-ban Admissions.

	Actual black admission rate	Hypothetical black admission rate under white coefficients	Average conditional black–white admission rate difference (col. 1 minus col. 2)
	(%)	(%)	(pp)
	(1)	(2)	(3)
A. Pre-ban			
Berkeley	56.7	5.6	51.1
	[43.6, 69.5]	[1.2, 11.4]	[38.7, 62.5]
UCLA	64.5	10.4	54.1
	[46.7, 80.6]	[2.2, 21.0]	[37.0, 70.5]
B. Post-ban			
Berkeley	31.3	13.5	17.8
	[20.4, 43.4]	[7.1, 20.6]	[9.3, 27.0]
UCLA	40.7	21.1	19.6
	[23.1, 60.0]	[7.9, 37.6]	[6.2, 34.1]

Notes — Each cell reports an estimate of either a black admission rate or a black–white admission rate difference using the EALS dataset. Ninety-five percent confidence intervals are computed using one thousand bootstrapped samples of each school-time period and are listed in brackets. Only black and white applications are used. Column 1 lists the actual black admission rate in the specified school-time period. Column 2 reports the black admission rate that is predicted to have prevailed if black applicants had been subjected to observed white admission standards, calculated by estimating a probit regression of admission on LSAT, GPA, inferred strength, a black indicator, and year fixed effects and then using the coefficients other than on the black indicator to predict admission probabilities for each applicant and accounting for the minor space-opening effect of a decline in black admission rates. Reported estimates are means of these predicted admission probabilities. Column 3 equals the difference between columns 2 and 1 and is an estimate of the average black-white admission rate difference for this school-time period's black applicants, conditional on covariates.

rate would have fallen to 8.0% had both black and white applicants been subjected to the same observed pre-ban white admission standards. Thus in spite of the large effects of the ban, the ban far from eliminated cross-sectional black admission advantages: holding the applicant pool constant at pre-ban levels, post-ban UC schools sustained average black admission advantages over observably similar whites equal to 22.5 percentage points (= 30.5% − 8.0%). This is consistent with admission offices either having shifted admission weight to non-racial black-correlates like family income and diversity essays in post-ban admissions (see Section 2.2), or with UC admission offices having placed uniquely large admission weight on non-racial black correlates even before the ban and relative to other schools (since I control for inferred strength).

5. Conclusion

Affirmative action bans can reduce black enrollment not only by inducing reductions in black admission advantages (demand contractions) but also by inducing reductions in applications (supply contractions) from black students who can still gain admission but prefer alternative schools that still practice affirmative action. I analyzed the case of Berkeley and UCLA law schools, which experienced severe declines in black applications, acceptances, and enrollment after the UC affirmative action ban even as black admission rates rose relative to whites at Berkeley. Data on a large sample of UC and non-UC applications as well as on their admission decisions made possible a unique joint analysis of supply and demand responses. I found that black attrition from UC applicant pools was driven mostly or entirely by less-credentialed black applicants who could no longer expect admission, yielding stronger post-ban black applicant pools. After holding applicant characteristics constant at pre-ban levels, I estimated that the ban cut black admission rates at both schools in half.

The results imply that even if supply responses had been muted—as might happen under a nationwide ban that eliminates affirmative-action-practicing alternatives—UC black enrollment would likely still have plummeted. Economically, the demand response is consistent with schools using non-racial admission factors that only partially sustained black admission advantages, in favor of sustaining other admission objectives: the selection-corrected post-ban black admission rate (31%) remained well above the rate that would prevail under observed white admission standards (8%) but was still only half the pre-ban black admission rate (61%). The supply response is consistent with black students still wanting to attend UC schools despite lower campus racial diversity, but choosing not to apply if they can no longer get in.

Effects may be different under a nationwide affirmative action ban. Notably, enrollment changes at less-elite schools under a nationwide ban may differ from the UC's experience depending on the cascading behavior of black students who no longer attend elite schools (Arcidiacono,2005; Epple et al.,2008). Less-elite black enrollment would be expected to decline if these new non-elite-attending black students abandon law school altogether, while it can actually increase if they are willing to trade down to lower-ranked schools where they can gain admission even without affirmative action. Hinrichs (2012) finds no effect of affirmative action bans on less-elite undergraduate minority enrollment though with sizeable standard errors; precise estimates of cascading behavior across hierarchies of undergraduate and professional schools are a priority for future work.

Finally, the results may bear on judicial debates. The Supreme Court has decided that affirmative action is unconstitutional whenever there are "workable race-neutral [non-racial] alternatives to achieve the diversity the university seeks" (*Grutter v. Bollinger* 2003).[33] Workability "does not require exhaustion of every conceivable race-neutral alternative…[nor] a university to choose between maintaining a reputation for excellence or fulfilling a commitment to provide educational opportunities to members of all racial groups" but does "require serious, good faith consideration of workable race-neutral alternatives" (*Grutter*). I find that UC law schools collected and used non-racial alternatives like family income and diversity essays, yet did not use them aggressively enough to keep black admission advantages from plummeting. This indicates by revealed preference that non-racial alternatives are far from workable from

33 "Race-neutral alternatives" include the use of non-racial black correlates like family income in admissions.

these elite law schools' perspectives, potentially bearing on courts' own legal judgments of workability.[34]

Appendix A. Models of demand and supply under an affirmative action ban

The first part of this appendix presents a simplified version of earlier models of admissions under an affirmative action ban (Chan and Eyster, 2003; Fryer et al., 2007; Epple et al., 2008) in order to show how a ban can induce a reduction in black admission advantages (a contraction in demand for black students) holding the applicant pool constant. The reduction in black admission advantages is large when legal non-racial admission factors like family income correlate only weakly with race (i.e. when the ban substantially raises the opportunity cost of admitting black students) and when schools are unwilling to sacrifice other objectives in order to sustain costlier racial diversity (i.e. when substitution effects dominate income effects).

The second part reproduces Card and Krueger (2004)'s model of application decisions in order to show how a ban can reduce black applications (a contraction in supply of black students). The reduction in black applications from marginal black candidates is large when demand contractions are large. The reduction in black applications from black students of all credential levels is large when black students' utility of attending an affected school substantially declines independently of their individual credential levels—e.g. via lower black campus representation following declines in other black students' applications, acceptances, and enrollment.

A.1. Demand contraction

Consider a school with concave preferences over the number of black enrollees $\bar{r}$ (short for "racial diversity") and the aggregate non-racial strength of enrollees $\bar{q}$ (short for "qualifications"). Each applicant is either black or white, the applicant pool is the same pre-ban and post-ban (abstracting here from supply effects), and all admitted students enroll. The school maximizes utility subject to a binding capacity constraint (it can admit no more than a fixed number $\bar{N}$ of applicants and must reject some applicants):

$$\max_{\bar{r},\bar{q}} u(\bar{r},\bar{q}) \quad \text{s.t.} \quad N(\bar{r},\bar{q}) \leq \bar{N}$$

where $N(\bar{r},\bar{q})$ is the minimum number of applicants that must be admitted in order to deliver $\bar{r}$ black admits and $\bar{q}$ aggregate non-racial strength. $N(\bar{r},\bar{q})$ is an implicit function of the joint distribution of race and non-racial strength in the applicant pool. The school faces a tradeoff in that the admission rule that maximizes the number of black admits is not the one that maximizes aggregate non-racial strength.

The school can admit applicants i on the basis of two pieces of information: non-racial strength q_i and a binary signal $BLACKSIGNAL_i \in \{0,1\}$ of black status. The black signal may be perfect (all black-signaled applicants are black and all white-signaled applicants are white) or diluted (not all black-signaled applicants are black and some white-signaled applicants are black). When the signal is diluted, I assume that dilution is orthogonal to non-racial strength. The optimal admission rule can then be characterized as a "rank-and-yank" rule that admits the $\bar{N}$ applicants that have highest rank according to:

$$RANK_i = q_i + \lambda BLACKSIGNAL_i$$

where λ is chosen to maximize utility. This is true because for any number of admitted black-signaled applicants, the school maximizes aggregate non-racial strength by adopting a threshold rule within each black signal value: only black-signaled applicants with non-racial strength above some $q^*_{BLACKSIGNAL=1}$ and white-signaled applicants with non-racial strength above some $q^*_{BLACKSIGNAL=0}$ are admitted. Rank-and-yank implements any such pair of threshold rules by setting weight λ equal to $q^*_{BLACKSIGNAL=0} - q^*_{BLACKSIGNAL=1}$.

When affirmative action is not banned, the black signal is perfect. Online Appendix Fig. 6a illustrates a feasible pair of optimal admission thresholds and illustrates its consequences for black and white applicants. To define the no-affirmative-action benchmark, let q^* be the level of non-racial strength above which there are exactly $\bar{N}$ applicants. This is the race-neutral threshold that would maximize aggregate non-racial strength and corresponds to a rank-and-yank admission rule with $\lambda = 0$. A school practicing affirmative action chooses $\lambda > 0$ and thus adopts a threshold admission rule for blacks at $q^*_{BLACKSIGNAL=1}$ and a separate higher threshold for whites at $q^*_{BLACKSIGNAL=0} > q^*_{BLACKSIGNAL=1}$. Relative to the no-affirmative-action benchmark, the school practicing affirmative action admits extra blacks (the grid fill pattern) and rejects extra whites (the solid fill pattern).

Online Appendix Fig. 6c illustrates an affirmative action budget set in $(\bar{r},\bar{q})$ space for the simple case of uniform distributions of non-racial strength within each race. The range of weights $\lambda \in [0,\infty)$ traces out the affirmative action ("AA") budget constraint. Point A is a potentially optimal bundle under affirmative action. The budget constraint is strictly convex because the first black applicant admitted through affirmative action is almost as strong as the white applicant that is rejected in order to make room. After that, stronger and stronger white applicants are rejected in order to make room for weaker and weaker black applicants.

An affirmative action ban prohibits the school from using a pure signal of race but allows it to use non-racial black-correlates like low family income that have plausible non-racial justification. I model this as dilution of the black signal with fraction p_{black} of black applicants and fraction p_{white} of white applicants signaled as black (e.g. those having family income below some threshold), with $p_{black} - p_{white} < 1$ and for simplicity $p_{black}, p_{white} \perp q_i$. The school increasing racial diversity above the no-affirmative-action benchmark now makes "mistakes" in the sense that the school rejects some applicants that have higher non-racial strength than accepted applicants of the same race, as illustrated in Online Appendix Fig. 6b. Thus an affirmative action ban raises the opportunity cost of admitting black applicants.

In the analytically tractable case of uniform distributions of non-racial strength within race,[35] the diluted black signal under an affirmative action ban ("BAN") raises the marginal rate of transformation of admitted blacks for non-racial strength by a factor that is decreasing in the purity of the black signal:

$$\frac{MRT^{BAN}_{\bar{r},\bar{q}}}{MRT^{AA}_{\bar{r},\bar{q}}} = \frac{1}{(p_{black} - p_{white})^2} > 1$$

The larger opportunity cost puts the affirmative-action-ban budget set inside the affirmative action budget set, illustrated in Online Appendix Fig. 6c. If substitution effects dominate, the school may respond to a ban by substantially contracting demand for black students (e.g. moving to bundle B)—sustaining the non-racial strength of admitted cohorts at the expense of racial diversity. But if income effects dominate, the school may barely contract demand for black

[34] The ongoing Supreme Court affirmative action case *Fisher v. Texas* centers on this question, in the context of Texas undergraduate admissions. The Court in 2012 returned *Fisher* to lower courts for strict scrutiny of whether a workable non-racial alternative exists. The Court heard new oral arguments on *Fisher* in December 2015.

[35] Without this or a similar assumption, the budget set can be non-convex over some intervals.

students (e.g. moving to bundle C) or even increase demand (if preferences are Giffen)—sustaining racial diversity at the expense of non-racial strength. Thus the degree to which a ban reduces black enrollment depends on the degree to which a ban dilutes the usable signal of race in admissions and on the substitutability of racial diversity for non-racial strength in the school's preferences.[36]

A.2. Supply contraction

Suppose a student has utility U_s of attending a school s, a probability P_s of gaining admission to s conditional on applying, and von Neumann–Morgenstern expected utility. Assume for simplicity that the student's admission decisions across different schools are independent conditional on P_s. Assume that there is a positive utility cost d of applying to each school, and let U_0 denote the utility of not attending any school. Let C denote the student's application set, comprising an ordered list of S schools with $U_1 \leq U_2 \leq \ldots \leq U_S$.

Taking $\{U_s, P_s\}$ as given, a student applies to a given school k if and only if the admission probability P_k exceeds a k-specific threshold. Specifically, let $C(\sim k)$ denote the optimal choice set when excluding school k from consideration, with $S(\sim k)$ denoting the number of schools in this set. The student will include k in her final choice set if and only if the expected value of applying to k exceeds the application cost:

$$P_k \left\{ \prod_{s=0}^{S(\sim k)} \pi_s \max\left[0, U_k - U_s\right] \right\} > d$$

where $\pi_s = P_s \cdot \Pi_{j=s+1}^{S} (1 - P_j)$ equals the probability that school $s \in C$ is the highest-utility school she gains admission to and $\pi_0 = \Pi_{j=1}^{S} (1 - P_j)$ equals the probability that the student is admitted to no school.[37] The expected value of applying to school k equals the probability of being admitted to k, times the probability that k is the best school she is admitted to.

An affirmative action ban can cause a black student to remove an affected school k from her application set through two channels.[38] First, the ban can reduce her probability of gaining admission P_k. In relatively formulaic contexts like law school admissions where students with sufficiently high test scores and grades are virtually guaranteed admission, the reduction in P_k may be zero for highly credentialed black students but large for less credentialed black students, inducing selective attrition from k's applicant pool. Second and to the extent that the student values a larger black student presence on campus, the ban can reduce utility U_k—via the lower likelihood of other black students applying, gaining admission, and enrolling—and thereby induce the student to remove k from her application set, especially if there are comparable unaffected schools that she can add to her application set. This substitution force (replacing affected schools with unaffected schools) can be arbitrarily large following even a small reduction in black admission advantages and can dissuade applications from black students uniformly across the credential distribution. The substitution force can be arbitrarily small when all schools are affected by the affirmative action ban.

36 Chan and Eyster (2003) adopt preference and technology restrictions to predict that the post-ban school introduces idiosyncratic noise—an imperfect black signal when black applicants are concentrated at lower levels of the non-racial strength distribution—to admission decisions. Fryer et al. (2007) analyze the case in which the post-ban school uses non-racial black correlates aggressively enough to admit the same number of black applicants as it did pre-ban.

37 She of course does not apply to any schools dominated by the no-school option.

38 Not modeled here is the possibility that an affirmative action ban raises a black student's likelihood of applying to an affected school, for example due to potentially higher signal value from gaining admission without affirmative action.

Appendix B. Details of EALS data coding

The first application year's LSAT scores are in a more compact scale than all other years', and I convert them to the modern scale using percentile rank. I de-mean GPA by year to account for modest grade inflation over time. I code "Hispanic", "Chicano/Mexican-American", and "Puerto Rican" as Hispanic. Undergraduate major is available in some years' raw data; it has low statistical power in subsamples and its use would limit the years available for analysis so I omit it. The admission decision for a small percentage of accepted students is classified as rejected when the applicant in fact accepted and deferred an admission offer. The relatively minor importance of this measurement error is suggested visually in Fig. 2b, where actual admission rates are close to 100% at high levels of academic strength, rather than plateauing at a smaller number. Year of college graduation is available in all years; I omit it from the analysis for simplicity but every qualitative result holds when also controlling for graduation year. The only other information in the raw data are indicators for whether the applicant took the LSAT more than once, whether the applicant withdrew an application before an admission decision was made, and whether the applicant accepted an admission offer. I exclude withdrawn applications from the analysis, and I do not have sufficient power to analyze matriculation decisions.

The raw data do not contain student identifiers, so for each year I create student identifiers by treating as coming from the same student those applications that match on all of the application-invariant variables. This is a powerful method for identifying applications submitted by the same student in largest part because GPA is coded to two decimal places. I exclude the fewer than one percent of observations for which this method implies that a single student submitted multiple applications to the same school.

Finally, I do not include the University of Michigan in the group of fifteen most-applied-to schools because it was subject to an affirmative action ban during the sample. I do not analyze Michigan as a treatment school because its bans were effective during the sample only in 2001 and 2006 and I do not have sufficient power to conduct year-by-year difference-in-differences. UC law schools at Davis and Hastings as well as public Texas law schools received few applications in the EALS and similarly do not permit robust inference.

Appendix C. Supplementary exhibits

Supplementary exhibits to this article can be found at http://dx.doi.org/10.1016/j.jpubeco.2016.02.006.

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