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THE EFFECTS OF AFFIRMATIVE ACTION BANS ON COLLEGE ENROLLMENT, EDUCATIONAL ATTAINMENT, AND THE DEMOGRAPHIC COMPOSITION OF UNIVERSITIES

Peter Hinrichs*

Abstract—I estimate the effects of affirmative action bans on college enrollment, educational attainment, and college demographic composition by exploiting time and state variation in bans. I find that bans have no effect on the typical student and the typical college, but they decrease underrepresented minority enrollment and increase white enrollment at selective colleges. In addition, I use the case study methods of Abadie and Gardeazabal (2003) and Abadie, Diamond, and Hainmueller (2010) and find that the affirmative action ban in California shifted underrepresented minority students from more selective campuses to less selective ones at the University of California.

I. Introduction

AFFIRMATIVE action in college admissions is one of today's most contentious social policy issues. Its supporters view it as a just response to past or present discrimination, stress the social benefits of producing minority role models and leaders, and claim that there are educational benefits to diversity. Its opponents contend that it is an impediment to achieving a race-blind society and may even be harmful to those it is intended to directly benefit. The issue has been in the headlines as affirmative action has been limited in various states in recent years by voters, courts, and elected officials. Most recently, voters in Michigan, Nebraska, and Arizona passed bans in 2006, 2008, and 2010, respectively, and the New Hampshire state legislature enacted a law banning affirmative action that went into effect on January 1, 2012. The U.S. Supreme Court is set to hear the affirmative action case *Fisher v. University of Texas* in its October 2012 term, and voters in Oklahoma will decide on the issue in November 2012.

This paper addresses the question of how affirmative action bans affect college enrollment, educational attainment, and the demographic composition of universities. If affirmative action raises the probability of admission for minorities at particular universities, then it is plausible that eliminating it will reduce minority enrollment. However, several factors may either magnify or diminish the impact of what happens at the admissions stage. First, eliminating affirmative action may have an effect on the behavior of potential students at the application or the enrollment stage. Probabilities of being admitted should affect the number and mix of

colleges a student applies to, and, moreover, an affirmative action ban may make minorities feel unwelcome and deter them from attending. Conversely, some individuals may be inclined to attend a university where it is known that race played no role in their admission decision. It may even be the case that an affirmative action ban increases the signaling value to underrepresented minorities of attending selective colleges. Second, universities may respond to an affirmative action ban by implementing policies that lessen its impact. For example, they may conduct greater outreach, decide to admit a larger number of students, or place greater weight on high school class rank in admissions. Third, even if an affirmative action ban reduces enrollment at a particular selective university, it is not clear what happens to those who are crowded out. Do they attend another selective university, do they "cascade down" to less selective institutions, or do they prefer to attend no college at all rather than attend their second choice? With all these issues in mind, it is not clear what the effect of an affirmative action ban on overall enrollment and the distribution of enrollment across schools actually is.

This paper uses information on which states have affirmative action bans in place in which years, along with data on college enrollment from the Current Population Survey (CPS), educational attainment from the American Community Survey (ACS), and college racial composition from the National Center for Education Statistics' Integrated Postsecondary Education Data System (IPEDS), in order to estimate the effects of affirmative action bans on college enrollment, educational attainment, and the demographic composition of various types of colleges. This paper differs from previous research in that I take a cross-state approach and estimate the effects of bans on the actual enrollment decisions of a random sample of the population.¹ I estimate difference-in-differences models that exploit the variation in bans across states and time. I find that the bans appear to have no effect on the typical student and the typical college, but they decrease underrepresented minority enrollment and increase white enrollment at selective colleges. Moreover, I use the case study methods of Abadie and Gardeazabal (2003) and Abadie, Diamond, and Hainmueller (2010) and find that the affirmative action ban in California shifted students from more selective campuses to less selective ones at the University of California. The picture that emerges from viewing the results from the three data sources together is that affirmative action bans do not affect who goes to college, but they have some effect on where people go to college.

¹ This is in contrast to the research by Card and Krueger (2005) discussed below, for example, as they focus on those who take the SAT.

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The rest of this paper is organized in the following manner: Section II places the paper in the context of previous research, section III contains the analysis of college attendance and educational attainment, section IV discusses the regression estimates of the effects of the bans on college racial composition, section V discusses the case studies, and section VI concludes.

II. Relation to Previous Research

Conceptually, an individual's college choice has four stages: participation, application, admissions, and enrollment. The participation stage involves the decision of whether to apply for college. If an individual decides to participate, the next stage is the choice of where to apply. Decisions in the third stage are in the hands of admissions committees and involve choosing which students are admitted among those who apply. In the fourth stage, students make choices about which college, if any, to attend among those to which they have been admitted. Affirmative action bans affect the third stage directly and may also have indirect effects at other stages.

Several recent studies have examined how affirmative action bans in California and Texas have affected decisions at the first stage or at the second stage conditional on having reached that stage. Dickson (2006), using a panel of Texas high schools, finds that the percentage of blacks and Hispanics who took the SAT fell when affirmative action was banned and did not recover when Texas implemented a policy to admit those in the top 10% of their high school class to any public university in the state.² Moreover, a lower percentage of whites took the test under the 10% plan but not after the initial ban on affirmative action. Card and Krueger (2005) use data on SAT takers in California and Texas to estimate how score-sending behavior of minorities changed relative to nonminorities over the time period that affirmative action bans went into effect in those states. They find little impact of affirmative action bans on where students send their scores, which suggests that affirmative action does not have an effect on the second stage of the college choice process for those who have reached that stage. In contrast, Long (2004a) finds that the gap between underrepresented minorities and others in sending SAT scores to top quintile colleges widens in California relative to control states when affirmative action was discontinued in California, although he does not find a statistically significant effect for Texas.³ Although both Card and Krueger (2005) and Long (2004a) make use

of high-quality individual-level data to take a focused look at one of the stages of the college choice process, they do share some common limitations. First, the samples in both papers are limited to those who take the SAT, so the results may be misleading if an affirmative action ban also affects whether people take the SAT. Second, although the decision of where to apply and send test scores is a stage in the college choice process, SAT-sending behavior is not the ultimate outcome of interest for policy.

Another set of papers uses data from Texas to focus on the third and fourth stages of the college choice. Tienda et al. (2003) show that the odds of admission among applicants fell at the University of Texas at Austin and Texas A&M for minorities relative to whites after affirmative action was banned.⁴ They also generally find a negative impact on minority enrollment relative to white enrollment among those who were admitted to the two universities. But it is worth bearing in mind that the enrollment effects on the population as a whole may differ from the enrollment effects conditional on being admitted. Kain, O'Brien, and Jargowsky (2005) find that among underrepresented minorities in Texas who attend a public institution of higher education within the state, the affirmative action ban had a negative effect on the probability of enrolling in selective institutions; moreover, this effect was not reversed by the Top 10% plan. Bucks (2005) also analyzes college choice among high school students from Texas and finds a lower probability among underrepresented minorities and a higher probability among others of enrolling in selective in-state public institutions in the post-affirmative action period. Neither Kain et al. (2005) nor Bucks (2005) is able to determine what happens to those who do not attend a public college in Texas. This is a limitation because there would be different implications in the case where an affirmative action ban at public universities causes people to attend out-of-state or private universities than in the case where it deters them from attending any college.

Arcidiacono (2005) uses data from the National Longitudinal Study of the Class of 1972 to estimate a structural model of college choice.⁵ Simulations suggest that removing affirmative action in admissions would have a modest negative impact on black college enrollment rates but a sizable negative impact on blacks attending colleges with high average SAT scores.

This paper estimates the effects of affirmative action at the enrollment stage. It differs from most previous research in several respects. First, I estimate the effects of affirmative action bans on a random sample of the college-aged population rather than limiting the sample to those who have taken the SAT, applied to a particular college, or chosen to attend a public university in a particular state. Second, I estimate the effects on actual enrollment decisions and educational attainment rather than on SAT sending. Third, I take a broad look at all states rather than focus-

² Also see Andrews (2007).

³ There are several differences in methodology between Card and Krueger (2005) and Long (2004a) that could lead to differences in results. First, Card and Krueger examine applications to a small number of particular schools, whereas Long examines application to schools broadly defined by quintile of standardized test scores. Second, Card and Krueger use a dummy for whether someone applies to a college as their main left-hand-side variable, and Long uses a count of the number of colleges an individual applies to. Third, Card and Krueger use data for more years (1994–2001) than Long does (1996 and 1999). Fourth, Card and Krueger use a difference-in-differences approach based on race and year, but Long incorporates a third dimension by using other states as controls.

⁴ See also Long and Tienda (2008).

⁵ Other recent structural papers are Hickman (2009) and Howell (2010).

TABLE 1.—SUMMARY STATISTICS OF CPS AND ACS DATA

Variable	Overall (1)	White (2)	URM (3)
<i>CPS sample (1995–2003)</i>			
Attends any college	0.398	0.459	0.268
Attends public college	0.316	0.358	0.228
Attends four-year college	0.258	0.314	0.140
Attends four-year public college	0.185	0.222	0.106
Male	0.509	0.500	0.502
N	13,488	9,750	3,738
<i>ACS sample (2005–2007)</i>			
Has attended college	0.602	0.641	0.477
Has associate degree	0.083	0.089	0.067
Has bachelor's degree or higher	0.231	0.264	0.123
Male	0.502	0.505	0.491
N	659,824	495,457	147,034

The table shows weighted means of variables and sample size by race and ethnicity for the Current Population Survey data and the American Community Survey data. All variables are binary. In the URM (underrepresented minority) category are blacks, Hispanics, and Native Americans.

ing on one particular state.⁶ A limitation of my estimates of the effect of bans on college attendance and educational attainment is the lack of information on the particular college attended. However, in the analysis on the racial composition of colleges, I am able to observe the racial composition of each college and estimate how various types of colleges are affected by affirmative action bans.

III. College Enrollment and Educational Attainment

A. Data and Empirical Methods

The data used in this section come from Current Population Survey (CPS) October School Enrollment Supplement files and from the American Community Survey (ACS). My sample is a random, nationwide sample of whites, blacks, Hispanics, and Native Americans who were 18 years old between 1995 and 2003. The CPS sample pools cross-sections of 18 year olds, while the ACS sample retrospectively uses 2005–2007 data for those who were 18 years old at some point between 1995 and 2003. The strength of the CPS is that it provides a more accurate measure of the relevant state, whereas the strength of the ACS is the large sample size.

I pool the October CPS for each year between 1995 and 2003 to estimate the relationship between affirmative action bans and contemporaneous school enrollment of 18 year olds. The CPS data allow me to determine whether someone attends college and, if so, whether that college is public or private and whether it is a two-year or four-year college. A useful feature of the CPS is that college students who are dependents of their parents are coded as being from the state where their parents live; thus, I am able to examine the effects of an affirmative action ban in the state in which an individual presumably resided while a senior in high school.⁷ Table 1 displays summary statistics for the CPS data.

⁶ A potential problem with the cross-state approach is that it assumes the absence of spillovers across states.

⁷ This assumes that parents do not move from one state to another after their child graduates from high school and also that the children themselves do not establish residency in another state.

I pool the ACS for 2005, 2006, and 2007. The ACS public use file for each year is a 1% random sample of the U.S. population. I use these data to retrospectively estimate the relationship between whether an affirmative action ban is in place at age 18 and educational attainment. As current state of residence may be an outcome of affirmative action bans and state at age 18 is unavailable, I link the data on affirmative action bans to the individual-level ACS data using state of birth. Thus, there may be some mismatch between what I take to be the relevant state for a person and what the relevant state actually is.⁸ Table 1 also displays summary statistics for the ACS data. It is noteworthy that a substantially larger proportion of those in the ACS have attended college than are currently attending college in the CPS data. This may be because some people in the CPS data set are still in high school, and it may also indicate that many people who eventually attend college do not begin immediately after finishing high school. In any case, the two data sets measure different concepts: the CPS contemporaneously measures college attendance at the current time, while the ACS retrospectively examines college attendance at any time before the survey is taken. The effects of an affirmative action ban on the two variables may be similar, or they may differ. For example, if the only result of an affirmative action ban at age 18 is to cause people to postpone college enrollment until age 19, then this would show up as no effect in the ACS data but a negative effect in the CPS data. If the only result of an affirmative action ban at age 18 is to prevent those who would have otherwise begun attending college at age 19 from attending college at all, then this would show up as no effect in the CPS data but a negative effect in the ACS data.

Table 2 shows which states have affirmative action bans in place for the fall admissions cycle for each year between 1995 and 2003. California, Florida, Texas, and Washington had an affirmative action ban in place at some point in this time period.⁹ My coding is consistent with previous studies and is based on the year the flagship public university in a state ended affirmative action in admissions.¹⁰ I drop observations from five states (Alabama, Georgia, Louisiana, Michigan, and Mississippi) that are in jurisdictions where there was important affirmative action litigation but that did not have outright bans on affirmative action.¹¹ It is also

⁸ There may also be some mismeasurement in what year someone turns 18 years old. The ACS is conducted throughout the year, and the public use data do not contain information on the month the survey is taken. Thus, the age and quarter of birth variables in the data are insufficient to recover the year someone turns 18. I assign the year at age 18 through the formula $yearat18 = yearofsurvey - age + 18$. Moreover, even if the year at age 18 were measured correctly, it is not necessarily the relevant year because some people finish high school and begin college earlier or later than age 18.

⁹ The bans in each of these states apply to public universities. Texas's ban also extended to private universities.

¹⁰ There is evidence that Florida State University began its affirmative action ban one year before the University of Florida did and also that Texas A&M began its affirmative action ban one year before the University of Texas did. The general results are robust to the alternate coding.

¹¹ Including these five states in the analysis does not cause any substantive changes in the results.

TABLE 2.—STATES WITH AFFIRMATIVE ACTION BANS AND PERCENTAGE PLANS

	1995	1996	1997	1998	1999	2000	2001	2002	2003
<i>States with affirmative action bans</i>									
California				X	X	X	X	X	X
Florida							X	X	X
Texas			X	X	X	X	X	X	X
Washington					X	X	X	X	X
<i>States with percentage plans</i>									
California							X	X	X
Florida						X	X	X	X
Texas				X	X	X	X	X	X

Texas A&M discontinued affirmative action in 1996, and Florida State University did so in 2000.

noteworthy that California, Florida, and Texas all implemented policies whereby achieving a certain high school class rank guaranteed acceptance at public universities. Table 2 also shows which states have “percentage plans” in which years. It is difficult to disentangle the effects of affirmative action bans and these percentage plans, partly because the two are highly collinear and partly because the percentage plans are potentially an outcome of banning affirmative action.¹² And if they are in fact an effect of banning affirmative action, then their effects should be attributed to affirmative action when estimating the reduced-form effects of banning affirmative action on enrollment and educational attainment.¹³ In any case, previous research has found that they do not raise minority enrollment to preban levels.¹⁴ The appendix contains more information about states’ affirmative action policies and percentage plans.

I estimate difference-in-difference linear probability models that exploit variation over time and state in affirmative action bans. I estimate these models separately for whites (those who report being white as their single race and are not Hispanic) and underrepresented minorities (Hispanics and those who report being at least part black or Native American). The models are of the form

$$y_{ist} = \text{ban}_{st}\alpha + \mu_s + \delta_t + \text{male}_i\theta + \varepsilon_{ist}. \quad (1)$$

Here y_{ist} is an outcome for individual i from state s in year t , and ban_{st} is a dummy for state s having an affirmative action ban in place in year t . The parameter of interest is α , the effect of an affirmative action ban on the outcome. The model also includes a full set of state dummies and year dummies. The variable male_i is a male dummy and θ is its coefficient, and ε_{ist} is the error term. I also estimate

¹² Long (2004b) writes, “There have been no states that have implemented an x% program without first dropping their affirmative action policies. Thus, these x% programs appear to be devised, implicitly or explicitly, to ameliorate the adverse effects of the elimination of race-based preferences.”

¹³ The same argument could be made about increased recruiting efforts or scholarship programs such as the University of Texas at Austin’s Longhorn Scholars program and Texas A&M’s Century Scholars program, both of which are targeted toward students at disadvantaged high schools.

¹⁴ See Long (2004b) and Kain et al. (2005). One reason for this result is that schools are not completely segregated, and so fewer than x% of minorities are in the top x% of their high school class. Another reason is that those who are toward the top of their class would be admitted even in the absence of the policy.

augmented models that include state-specific linear time trends.

B. Results

The top part of table 3 explores the relationship between affirmative action bans and college attendance using CPS data. The table reports coefficient estimates with standard errors that are robust to clustering at the state level. The estimate of equation (1) reported in the first column of the first row suggests that an affirmative action ban is associated with a 1.64 percentage point higher rate of college attendance among underrepresented minorities. A scenario under which this could occur is if bans do not affect the decision to attend college among those who would have attended in the absence of the ban but result in outreach efforts that cause additional minorities to attend college. However, the estimate is imprecise and not statistically distinguishable from 0. Moreover, the estimate becomes negative and very small in column 2, which includes state-specific time trends.

Although affirmative action bans at public universities may not reduce the overall rate of college attendance among minorities, they may affect the type of college attended. For instance, they may cause a shift away from public colleges or from four-year colleges to two-year colleges. However, the remaining rows of table 3 for the CPS sample suggest that this is not the case, although here I cannot rule out the possibility that affirmative action bans shift minorities from more selective to less selective in-state public universities or to out-of-state public universities.¹⁵

Table 3 also provides an analysis of educational attainment using ACS data. The first row of the bottom part of the table, much like the first row of the table for CPS data, suggests that affirmative action bans do not have an effect on whether underrepresented minorities attend college. The second and third rows of the bottom part of the table estimate the effects of affirmative action bans on receiving an associate degree and on receiving a bachelor’s degree or

¹⁵ This raises the possibility of general equilibrium effects. For instance, minorities crowded out of a public university in their home state may decide to attend college in another state, thereby affecting the market for higher education in that other state. This would be a limitation for the across-state comparison of the difference-in-differences strategy. However, the results using IPEDS data presented later in table 7 do not support the view that affirmative action bans result in migration across state lines to attend college.

TABLE 3.—EFFECTS OF AFFIRMATIVE ACTION BANS ON COLLEGE ATTENDANCE AND EDUCATIONAL ATTAINMENT

Outcome	Underrepresented Minority		White	
	Without Time Trends (1)	With Time Trends (2)	Without Time Trends (3)	With Time Trends (4)
<i>CPS Sample</i>				
Attends any college	0.0164 [0.0269]	−0.0067 [0.0296]	−0.0168 [0.0215]	−0.0649 [0.0486]
Attends public college	0.0093 [0.0231]	−0.0050 [0.0250]	−0.0497 [0.0210]*	−0.0861 [0.0593]
Attends four-year college	0.0187 [0.0190]	0.0124 [0.0205]	0.0179 [0.0185]	−0.0348 [0.0453]
Attends four-year public college	0.0146 [0.0182]	0.0186 [0.0213]	−0.0079 [0.0196]	−0.0499 [0.0501]
<i>N</i>		3,738		9,750
<i>ACS Sample</i>				
Has attended college	0.0123 [0.0069]	0.0131 [0.0070]	−0.0065 [0.0030]*	−0.0023 [0.0055]
Has associate degree	0.0082 [0.0026]**	0.0051 [0.0022]*	0.0082 [0.0047]	−0.0006 [0.0023]
Has bachelor's degree or higher	−0.0010 [0.0037]	0.0033 [0.0032]	−0.0039 [0.0081]	−0.0101 [0.0123]
<i>N</i>		147,034		495,457

Each cell corresponds to a separate regression estimate of equation (1). Each row corresponds to an outcome. Columns 1 and 2 estimate the model for underrepresented minorities, while columns 3 and 4 estimate the model for whites. Columns 2 and 4 contain state-specific time trends, while columns 1 and 3 do not. The table displays estimates for the ban variable, with standard errors corrected for clustering at the state level in brackets. All models also include a gender dummy, year dummies, and state dummies. A single asterisk denotes significance at the 5% level, and a double asterisk denotes significance at the 1% level. In the under-represented minority category are blacks, Hispanics, and Native Americans.

higher. If an affirmative action ban shifts minorities away from four-year colleges, it may increase the proportion who have an associate degree. The estimates in the second row give some indication that this may be the case, although the magnitudes of the estimates are quite small. An affirmative action ban may lower the probability of receiving a bachelor's degree or higher for minorities if it displaces them from four-year colleges or shifts them to colleges that have lower graduation rates. Alternatively, an affirmative action ban may increase minority graduation rates if it reduces a mismatch between minorities and the type of college they attend. Nonetheless, the results in the third row of the bottom part of table 3 for the ACS sample do not lend much support to either of these possibilities.

The results in the third and fourth columns of table 3 suggest that banning affirmative action does not have a positive impact on white college attendance. Of the seven estimates without state-specific time trends, two are significant but in an unexpected direction. Moreover, these estimates do not maintain their significance when state time trends are included in the models.

IV. Racial Composition of Universities: Regression Estimates

A. Data and Empirical Methods

I use data from the National Center for Education Statistics' Integrated Postsecondary Education Data System (IPEDS) to estimate the effects of affirmative action bans on the racial composition of universities of various types. IPEDS has extremely broad coverage. It includes data from all institutions of higher education that have a Program Participation Agreement with the Department of Education to

provide student financial aid under federal programs such as the Federal Pell Grant Program, the Ford Direct Loan Program, and the Federal Perkins Loan Program.¹⁶ I use enrollment data by race for each year between 1995 and 2003.¹⁷ Table 4 reports summary statistics.

I estimate models of the form

$$y_{ist} = \text{ban}_{st}\alpha + \mu_i + \delta_t + \eta_{st} + \varepsilon_{ist}. \quad (2)$$

Here y_{ist} is the percentage of first-time degree-seeking undergraduates who are of a particular demographic group at university i in state s at time t , ban_{st} is a dummy for whether a ban is in effect in state s in year t , μ_i is a full set of institution dummies, δ_t is a full set of year dummies, η_{st} is a full set of state-specific linear time trends, ε_{ist} is a disturbance, and α is the parameter of interest. I report results for various subsamples of universities. The regressions are weighted by enrollment, and I report standard errors that are robust to clustering at the state level.

In order to determine whether a ban has a different effect on outcomes depending on the selectivity level of a college, I also estimate models of the form

$$y_{ist} = x_{is}\text{ban}_{st}\beta + \text{ban}_{st}\alpha + \mu_i + \delta_t + \eta_{st} + \varepsilon_{ist}, \quad (3)$$

where x_{is} is either the admission rate or the midpoint of the 25th to 75th percentile range of SAT scores (or SAT-

¹⁶ It may also include data on other institutions that submit data voluntarily.

¹⁷ I drop from the sample institutions that were missing data in at least one year, as well as historically black colleges and universities. I also drop a small number of colleges where either the total or within-gender enrollment numbers by racial group do not add up to overall enrollment. In regressions that are restricted to public universities, I include only institutions that are coded as being public in every year of the sample.

TABLE 4.—SUMMARY STATISTICS OF IPEDS DATA

Racial Group	Type of Institution					
	Four-Year (1)	Public Four-Year (2)	U.S. News Top Two Tiers (3)	Public U.S. News Top Two Tiers (4)	U.S. News Top 50 (5)	Public U.S. News Top 50 (6)
Asian	7.28	7.54	13.22	13.21	19.72	22.32
Black	7.02	6.96	5.72	5.76	6.05	5.79
Hispanic	7.16	7.35	6.53	6.36	7.00	7.38
Native American	0.83	0.98	0.53	0.52	0.53	0.51
White	75.54	75.74	71.49	72.63	63.22	62.15
N	9,108	3,393	891	450	387	126

The table displays weighted means of the percentage of students at the various types of institutions (columns) that belong to the various racial groups (rows).

TABLE 5.—EFFECT OF BANS ON REPRESENTATION BY RACE

Racial Group	Type of Institution					
	Four-Year (1)	Public Four-Year (2)	U.S. News Top Two Tiers (3)	Public U.S. News Top Two Tiers (4)	U.S. News Top 50 (5)	Public U.S. News Top 50 (6)
Asian	0.4502 [0.2728]	0.6286 [0.4365]	0.5520 [0.2144]*	0.9158 [0.4377]*	0.8650 [0.4527]	1.4298 [0.8011]
Black	-0.4069 [0.2364]	-0.5112 [0.3392]	-0.9964 [0.2337]**	-1.1778 [0.3006]**	-1.6424 [0.6586]*	-1.7429 [0.7985]
Hispanic	-0.6891 [0.4109]	-0.6808 [0.4835]	-1.0811 [0.5142]*	-1.2213 [0.6490]**	-1.8262 [0.5646]**	-2.0332 [0.6956]*
Native American	-0.0718 [0.0444]	-0.1072 [0.0603]	-0.1204 [0.0808]	-0.1397 [0.1010]	-0.3772 [0.1361]*	-0.4660 [0.1103]**
White	0.9612 [0.4402]*	0.8542 [0.3720]*	1.8258 [0.4266]**	1.7790 [0.6989]*	3.0897 [0.6705]**	2.9268 [0.9484]*
N	9,108	3,393	891	450	387	126

Each cell corresponds to a separate (weighted) regression estimate of equation (2). The left-hand side variable is the percentage of first-time degree-seeking undergraduates who are members of the various racial groups. Rows corresponds to a racial group, and columns correspond to types of institutions. The table displays coefficient estimates on the ban variable, with standard errors corrected for clustering at the state level in brackets. All models include a full set of year dummies, a full set of institution dummies, and a full set of state-specific time trends. A single asterisk denotes significance at the 5% level, and a double asterisk denotes significance at the 1% level.

equivalent ACT scores) reported in the 1995 *U.S. News & World Report* college rankings.¹⁸

B. Results

Table 5 gives estimates of the effects of affirmative action bans on the racial composition of various types of colleges. The overall effect for whites at four-year colleges is marginally significant, while the effects for other groups are not significant. This result is generally consistent with the results from section III that showed no effect on enrollment and graduation from four-year institutions. One explanation for these results is that the typical university is not selective and will thus not be directly affected by an affirmative action ban. An alternative explanation is that the typical university is not directly affected because the typical university is private and, with the exception of Texas, the affirmative action bans considered here apply only to public universities. Columns 2 to 6 provide some support for both of these possibilities. The negative effect of bans on underrepresented minority enrollment is larger in magnitude when the sample is limited to public universities, the 115 universities in the top two tiers of the 1995 *U.S. News &*

World Report college rankings, or the top 50 universities in the *U.S. News* rankings. Column 6 suggests that banning affirmative action at a public university in the top 50 of the *U.S. News* rankings is associated with a decrease in black enrollment of roughly 1.74 percentage points, a decrease in Hispanic enrollment of roughly 2.03 percentage points, and a decrease in Native American enrollment of roughly .47 percentage points. According to table 4, the shares of students at these universities who are black, Hispanic, and Native American are 5.79%, 7.38%, and .51%. Thus, the changes in representation caused by affirmative action bans are very large in relative terms. Banning affirmative action at these universities is also associated with an increase in white enrollment of roughly 2.93 percentage points and an increase in Asian enrollment of roughly 1.43 percentage points. In relative terms, these changes are not as large as the changes for underrepresented minorities due to the already high representation of whites and Asians at selective universities. In results not reported here, I also generally find that the results do not vary by gender.

As an additional test of whether more selective universities are more heavily affected by bans, table 6 reports results that interact the *ban* variable with the university's admission rate or SAT score for the schools in the top two tiers of the 1995 *U.S. News* rankings. Most of the coefficients here are significant. The negative effects of bans on

¹⁸ Two universities report only a single number rather than the 25th and 75th percentiles. In those cases, the single number that is reported is used.

TABLE 6.—INTERACTIONS OF BAN WITH ADMISSION RATE AND SAT SCORE FOR SCHOOLS IN TOP TWO TIERS OF *U.S. NEWS* RANKINGS

<i>Racial Group</i>	<i>Ban Interacted with Admission Rate</i>		<i>Ban Interacted with SAT Score</i>	
	Main Effect (1)	Interaction (2)	Main Effect (3)	Interaction (4)
Asian	6.6951 [1.0875]**	−0.0933 [0.0177]**	−17.1264 [1.3185]**	0.0161 [0.0012]**
Black	−3.1238 [0.3101]**	0.0323 [0.0012]**	3.4840 [0.6321]**	−0.0041 [0.0004]**
Hispanic	−6.7381 [1.3941]**	0.0860 [0.0291]**	11.2924 [6.1513]	−0.0113 [0.0054]*
Native American	−0.4847 [0.0958]**	0.0055 [0.0009]**	0.7595 [0.3056]*	−0.0008 [0.0002]**
White	3.5346 [1.5277]*	−0.0260 [0.0219]	4.3878 [4.3164]	−0.0023 [0.0036]
<i>N</i>	891		891	

The table shows (weighted) regression estimates of equation (3). Within each row, columns 1 and 2 correspond to one regression and columns 3 and 4 to another. The left-hand-side variable is the percentage of first-time degree-seeking undergraduates who are members of the various racial groups. Columns 1 and 2 report the estimated coefficients on the ban variable and on its interaction with the admission rate. Columns 3 and 4 report the estimated coefficients on the ban variable and on its interaction with the average of the 25th and 75th percentiles of the SAT score distribution at the institution. Standard errors corrected for clustering at the state level are in brackets. All models include a full set of year dummies, a full set of institution dummies, and a full set of state-specific time trends. A single asterisk denotes significance at the 5% level, and a double asterisk denotes significance at the 1% level.

TABLE 7.—STATE-LEVEL REGRESSIONS OF IN-STATE COLLEGE ATTENDANCE ON BAN DUMMY

	<i>Type of Institution</i>					
	Four-Year (1)	Public Four-Year (2)	<i>U.S. News</i> Top Two Tiers (3)	Public <i>U.S. News</i> Top Two Tiers (4)	<i>U.S. News</i> Top 50 (5)	Public <i>U.S. News</i> Top 50 (6)
Mean	75.45	85.03	67.35	80.61	62.32	86.06
Coefficient (SE)	1.5764 [1.2507]	1.4374 [1.1613]	6.3881 [3.0874]*	4.7769 [2.8683]	3.0780 [3.1383]	1.3171 [1.3654]
<i>N</i> (states)	306	306	306	306	306	306

Each column corresponds to a group of universities. The first row gives the percentage of college-attending students within that group who attend an in-state college. The second row reports results from estimating equation (4), a weighted state-level regression of this variable on the ban dummy. The table displays coefficient estimates on the ban variable, with standard errors corrected for clustering at the state level in brackets. All models include a full set of year dummies, a full set of institution dummies, and a full set of state-specific time trends. A single asterisk denotes significance at the 5% level.

underrepresented minority enrollment are larger the lower the admission rate or the higher the SAT score. For example, at a school with an admission rate of 60%, a ban is predicted to reduce black representation by 1.19 percentage points ($-3.1238 + .0323 \times 60$) and Hispanic representation by 1.58 percentage points ($-6.7381 + .0860 \times 60$). But at a school with an admission rate of 30%, a ban is predicted to reduce black representation by roughly 2.15 percentage points and Hispanic representation by roughly 4.16 percentage points. Additionally, these results suggest that there may be an increase in Asian enrollment when affirmative action is prohibited at a university that is sufficiently selective.

The IPEDS regressions capture students by state of college attendance rather than state of residence, so one concern is selective migration for college across state lines caused by affirmative action bans. For 1996, 1998, and 2000–2003, the IPEDS includes data on the distribution of state of residence for freshmen at each college. I can use these data to test, in a limited way, whether affirmative action bans result in movement across state lines for college. Table 7 reports results from a state-level regression of the percentage of students attending college who are attending a college in their home state on the ban dummy. That is, I estimate the model

$$y_{st} = \text{ban}_{st}\alpha + \mu_s + \delta_t + \eta_{st} + \varepsilon_{st}, \quad (4)$$

where y_{st} is the percentage of college-going freshmen from state s at time t who are attending college within their state

of residence, ban_{st} is a dummy for whether a ban is in effect in state s in year t , μ_s is a full set of state dummies, δ_t is a full set of year dummies, η_{st} is a full set of state-specific linear time trends, ε_{st} is a disturbance, and α is the parameter of interest. The results do not suggest migration to out-of-state colleges in response to affirmative action bans. However, the data here are limited by the fact that they are not disaggregated by racial group, which leaves open the possibility that there is no net effect but that there is out-migration for underrepresented minorities and immigration for whites and Asians.

V. Racial Composition of Universities: Case Studies

The regression estimates presented in section IV may obscure the fact that certain universities may be more affected by affirmative action bans than others. In this section, I evaluate this possibility by using the synthetic control method for comparative case studies developed in Abadie and Gardeazabal (2003) and Abadie et al. (2010). The setup of the synthetic control method is that there is a treatment that goes into effect at some point in one treatment unit but not in the pool of potential control units. The researcher specifies a set of variables to use in forming a control, and the synthetic control is chosen to be the convex combination of the potential control units that most closely matches the treatment unit in the value of these variables. For example, a

researcher may choose the synthetic control to match the value of an outcome variable in the treatment unit in the pre-treatment period. Projecting the value of the outcome variable in the synthetic control into the posttreatment period then gives an estimate of the counterfactual: what would have happened to the outcome variable in the treatment unit had the treatment not gone into effect. Taking the difference between the outcome variable in the treated unit and the synthetic control in the posttreatment period gives an estimate of the treatment effect. Under this method, the data play a role in choosing the control group rather than the control group being chosen solely by the researcher.

Following Abadie et al. (2010), the model is

$$y_{it} = \text{ban}_{it}\alpha_{1t} + \theta_i Z_i + \lambda_t \mu_t + \delta_t + \varepsilon_{it}. \quad (5)$$

This is essentially a standard difference-in-differences equation, although it allows a different treatment effect in each year. It also allows the “fixed” effect for the cross-section units to vary over time, albeit in a particular way. If the values of the variables used for matching are stacked into the vector X_1 for the treated unit and the matrix X_0 for the potential control units, then I choose the weights W used in forming the synthetic control to minimize

$\sqrt{(X_1 - X_0 W)' V (X_1 - X_0 W)}$. The matrix V weights the variables used in the matching. I choose V to minimize the mean squared prediction error over the entire pretreatment period. The estimates of the α_{1t} are the differences between the outcome in the treatment unit and the synthetic control unit for the posttreatment period. In other words, $\hat{\alpha}_{1t} = y_{1t} - \sum_{j=2}^{J+1} w_j y_{jt}$, where the treated unit is the first of the $J + 1$ total cross-section units and where w_j is the weight put on the j th cross-section unit in the synthetic control.

I use the synthetic control method to estimate the effects of the affirmative action ban in California on underrepresented minority enrollment at the campuses of the University of California. To gain power, I pool the eight campuses into two aggregates.¹⁹ The more selective aggregate consists of the four campuses that were the highest ranked of the eight in the 1995 *U.S. News & World Report* rankings: UC Berkeley, UCLA, UC Davis, and UC San Diego. The less selective aggregate consists of the other four campuses that offered undergraduate education in this time period: UC Irvine, UC Riverside, UC Santa Barbara, and UC Santa Cruz.²⁰ The

pool of potential control units for the more selective aggregate contains universities ranked between 1 and 47 of the 1995 *U.S. News & World Report* rankings, and the pool of potential control units for the less selective aggregate contains universities ranked between 44 and 115 in the rankings.²¹ I exclude from the donor pool universities that are located in the four ban states, as well as those located in the five states that were excluded from the regressions of sections III and IV. I use IPEDS data from 1986, 1988, and 1990 to 2003 in order to view enrollment trends over an extended period of time. I form the synthetic control by matching with the racial composition from each year in the pre-ban period, the fraction of those in the state in 1990 who are underrepresented minorities according to 1990 Census Summary Tape File 1, and per capita income in the state in 1995 according to the 1997 *Statistical Abstract of the United States*. I estimate the effects of a ban by comparing the racial composition in the treated unit to that in the synthetic control in the post-ban period.

Figure 1 shows the effect of the affirmative action ban at the more selective University of California campuses. The weights given to particular universities in forming the synthetic control and the predictor balance are shown in table 8. The ban has a clear negative effect on underrepresented minority enrollment, as there is a sharp drop in underrepresented minority enrollment at these campuses relative to their synthetic control in 1998. But interestingly, there is a slight rebound over time. The enrollment rebound could be due to the University of California's enacting policies to lessen the impact of the ban. Such policies include the policy of automatically admitting those in the top 4% of their high school class, greater outreach, or altered admissions or scholarship criteria. Figure 2 shows the effect at the less selective UC campuses. At these campuses, there is a rise in underrepresented minority representation when the affirmative action ban goes into effect, presumably from students cascading down to these campuses from the more selective ones. Overall, the synthetic control analysis shows that affirmative action bans can cause a large fall in underrepresented minority enrollment at certain universities and an increase at others. Estimating the effects only on the mean university may mask these distributional effects.

VI. Conclusion

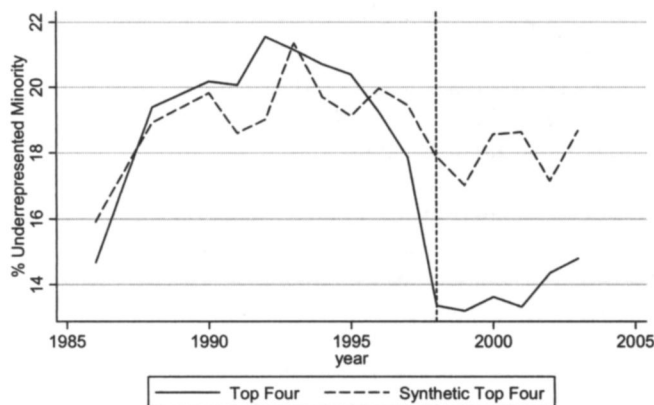
This paper finds that for the typical student and the typical college, affirmative action bans have no effect. However, affirmative action bans decrease underrepresented minority enrollment at selective colleges. For example, the results suggest that banning affirmative action at a public

¹⁹ Previous versions of this paper contained separate estimates for each campus, as well as estimates for the University of Texas at Austin, Texas A&M, the University of Florida, Florida State University, the University of Washington, Washington State University, Rice University, and Stanford University. The general results are similar to the results with the two aggregates: the effect of an affirmative action ban varies by university, and the bans cause a reshuffling of students from more selective to less selective institutions.

²⁰ UC Berkeley ranked 26th overall, UCLA ranked 28th, UC Davis ranked 40th, UC San Diego ranked 43rd, and UC Irvine ranked 48th. UC Riverside, UC Santa Barbara, and UC Santa Cruz were in the second tier, which consists of ranks 51 to 115.

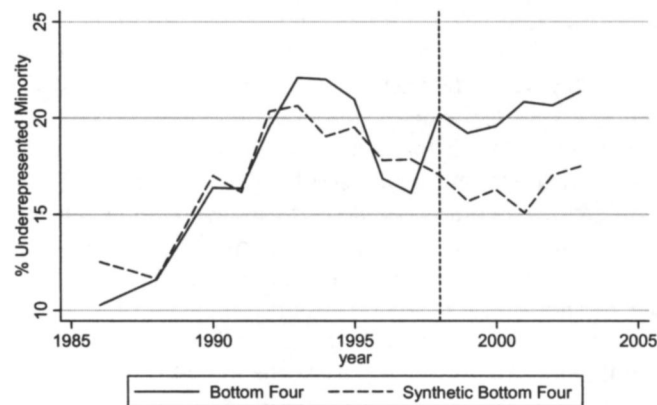
²¹ This means that the donor pool for the top four UC campuses includes institutions ranked higher than the fifth UC campus and that the donor pool for the bottom four UC campuses includes institutions ranked lower than the fourth UC campus. Thus, each of the synthetic controls is a combination of universities that are ranked similarly to the treated unit it serves as a control for. But the results are substantively unaffected when not imposing the restriction that the controls be ranked similarly to the treated unit.

FIGURE 1.—BAN AT TOP FOUR UC CAMPUSES



The four campuses are UC Berkeley, UCLA, UC Davis, and UC San Diego.

FIGURE 2.—BAN AT BOTTOM FOUR UC CAMPUSES



The four campuses are UC Irvine, UC, Riverside, UC Santa Barbara, and UC Santa Cruz.

TABLE 8.—WEIGHTS AND PREDICTOR BALANCE WITH SYNTHETIC CONTROL METHOD

A: Weights					
Top Four UC Schools			Bottom Four UC Schools		
University in Synthetic Control	Weight		University in Synthetic Control	Weight	
Rutgers–New Brunswick	.812		Fordham University	.499	
Harvard University	.161		Colorado School of Mines	.222	
Dartmouth College	.027		Polytechnic University	.138	
			George Washington University	.064	
			University of Massachusetts–Boston	.042	
			Stevens Institute of Technology	.031	
			Illinois Institute of Technology	.004	
B: Predictor Balance					
Matching Variable	Year	Top Four UC Schools		Bottom Four UC Schools	
		Treated	Control	Treated	Control
% URM at institution	1986	14.68	15.91	10.26	12.54
% URM at institution	1988	19.40	18.94	11.62	11.67
% URM at institution	1990	20.18	19.83	16.35	16.99
% URM at institution	1991	20.07	18.61	16.31	16.13
% URM at institution	1992	21.55	19.04	19.51	20.34
% URM at institution	1993	21.14	21.35	22.06	20.62
% URM at institution	1994	20.69	19.71	21.98	19.03
% URM at institution	1995	20.39	19.13	20.96	19.51
% URM at institution	1996	19.25	19.97	16.84	17.78
% URM at institution	1997	17.87	19.46	16.08	17.84
% URM in state	1990	33.48	23.40	33.48	24.04
Per capita income in state	1995	24,091	27,611	24,091	28,595

URM: "underrepresented minority," which includes blacks, Hispanics, and Native Americans. UC: University of California.

university in the top 50 of the *U.S. News* rankings is associated with a decrease in black enrollment of roughly 1.74 percentage points, a decrease in Hispanic enrollment of roughly 2.03 percentage points, and a decrease in Native American enrollment of roughly .47 percentage points. Since the bases are already small, these effects are very large in relative terms. Moreover, the results using the synthetic control method of Abadie and Gardeazabal (2003) and Abadie et al. (2010) confirm that affirmative action bans have an especially large effect at particular selective institutions. Thus, the results in this paper are consistent with Arcidiacono (2005), in which affirmative action bans affect the distribution of colleges attended but have only a small effect on overall college attendance rates. However, it

is worth bearing in mind that the states with affirmative action bans in the time period studied in this paper are home to both a large minority population and selective public institutions. Thus, the effect of an affirmative action ban in the typical state may not be quite as large as the effects found in this paper.²² But taking the results at face value, the practical importance of the findings hinges on three important issues: (1) whether college quality matters for later labor market outcomes, (2) whether underrepresented

²² To be sure, the effects of bans on enrollment as a share of baseline minority enrollment may be the same in high-minority and low-minority states, but this amounts to a larger absolute change in high-minority states.

minorities are “mismatched” at selective colleges, and (3) whether college diversity matters for later outcomes. Existing evidence on each of these issues is mixed.²³ But to the extent that college quality matters for later outcomes, the fact that there is a decline in underrepresented minority enrollment at selective institutions when affirmative action is banned would be detrimental to underrepresented minorities. On the other hand, to the extent that there is mismatch, the results in this paper may be welfare enhancing for all. To the extent that college diversity is important for later-life outcomes, the shift in minority enrollment from more selective to less selective colleges may be detrimental to students at more selective colleges but beneficial for students at less selective colleges.

The affirmative action debate will likely continue in the United States for years to come. As long as there remains racial inequality in income and educational attainment, affirmative action in college admissions will be viewed as a policy lever that can potentially help correct the imbalance. But it appears that the tide is beginning to turn against affirmative action. This has prompted concern that racial inequality in education may widen. This paper finds mixed evidence for that concern: the typical student and the average college seem to be unaffected by affirmative action bans, but particular selective institutions are heavily affected.

²³ On the first issue, see Black and Smith (2004), Dale and Krueger (2002), and Hoekstra (2009). On the second issue, see Kane (1998), Rothstein and Yoon (2009), and Sander (2004). On the third issue, see Arcidiacono and Vigdor (2010) and Hinrichs (2011). For a general review of the literature on affirmative action, see Holzer and Neumark (2000, 2006).

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APPENDIX

Affirmative Action Policies and Percentage Plans

Arizona

Arizona voters voted to ban affirmative action in 2010.

California

California’s ban went into effect in 1998 following an earlier decision by that state’s board of regents and after California voters passed proposi-

tion 209. The board of regents' decision has since been overturned, and now Proposition 209 is what holds the ban in place. Under California's Eligibility in the Local Context policy, those in the top 4% of their high school class are guaranteed admission to at least one campus of the University of California.

Florida

Florida's affirmative action ban is a result of then-Governor Jeb Bush's One Florida plan. Under Florida's Talented 20 Program, those in the top 20% of their high school class are guaranteed admission to at least one public university in Florida.

Georgia

The University of Georgia's affirmative action policy was struck down by a circuit court ruling, although opinions differ as to whether this banned affirmative action in that circuit (Hebel, 2001a, 2001b). Nonetheless, the University of Georgia eliminated affirmative action beginning in the fall 2002 admissions cycle (Hebel, 2001a, 2001b). Alabama is in the same circuit as Georgia and is also dropped from the sample. The other state in that circuit is Florida, which already had its own affirmative action ban.

Michigan

The University of Michigan made major revisions to its affirmative action policy to make it more flexible in the wake of the Supreme Court's 2003 *Gratz v. Bollinger* decision (University of Michigan News Service, 2003), but it did not eliminate affirmative action at that time. However, public universities in Michigan were not allowed to use affirmative action

in admissions as a result of Michigan voters' passing Proposal 2 in November 2006, although this initiative has been challenged in court.

Nebraska

Nebraska voters voted to ban affirmative action in 2008.

New Hampshire

The New Hampshire state legislature enacted a law banning affirmative action that went into effect on January 1, 2012.

Texas

Texas's affirmative action ban went into place as a result of a ruling by the Fifth Circuit Court of Appeals in the case of *Hopwood v. State of Texas*. This ruling was overturned by the Supreme Court's 2003 decisions in *Gratz v. Bollinger* and *Grutter v. Bollinger*, and universities in Texas are now permitted to use affirmative action. The University of Texas at Austin reintroduced affirmative action for its fall 2005 admissions cycle (University of Texas at Austin Office of Admissions, 2006). Under Texas's law HB 588, those in the top 10% of their high school class are guaranteed admission into any public university in Texas. Louisiana and Mississippi are part of the same circuit as Texas, in which affirmative action was outlawed as a result of the *Hopwood* ruling, but were under federal desegregation orders that pointed them in a conflicting direction (Healy, 1998).

Washington

Washington's affirmative action ban is due to voters passing Initiative 200 in 1998.