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ABSTRACT

A Cure for Discrimination? Affirmative Action and the Case of California Proposition 209*

Proposition 209, enacted in California in 1996 and made effective the following year, ended state affirmative action programs not only in education, but also for public employment and government contracting. This paper uses CPS data and triple difference techniques to take advantage of the natural experiment presented by this change in state law to gauge the labor market impacts of ending affirmative action programs. Employment among women and minorities dropped sharply, a change that was nearly completely explained by a decline in participation rather than by increases in unemployment. This decline suggests that either affirmative action programs in California had been inefficient or that they failed to create lasting change in prejudicial attitudes.

JEL Classification: J71, J78

Keywords: economics of gender and minorities, affirmative action, Proposition 209, discrimination

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1 Introduction

Introducing and removing affirmative action are not opposite sides of the same coin. Proponents of affirmative action maintain that it will provide a long-term cure for discrimination by allowing victims to demonstrate their skill and worth, thus changing prejudicial attitudes. Under this scenario, if affirmative action "works," then when it is time to get rid of the program there will be no deleterious effects for minorities. Opponents of these controversial programs, however, argue that it does not address the root source of inequality and, moreover, that it may create labor market inefficiencies and result in reverse discrimination against white males. Both sides, therefore, suggest that an effective affirmative action program would cause minority employment to rise, but they disagree on whether this increase is efficient and whether it would be sustainable if formal affirmative action were ended.

To date, there has been little opportunity to measure the impact of removing affirmative action programs. While federal support for enforcement has ebbed and flowed and Supreme Court rulings in the past decade have chipped away at affirmative action, it is difficult to say whether concurrent changes in minority outcomes were due to affirmative action policy or other trends in inequality. A similar problem plagued attempts to measure the impact of instituting affirmative action in earlier years. While minorities and women made gains in the labor market in the seventies and eighties, it is not clear what portion of this was due to affirmative action and what was the result of other influences. Empirical studies of the impact of affirmative action on labor markets have relied on differences in outcomes for government contractors, who are subject to the program, and non-contractors, who are not. While these studies have provided evidence of minority gains among contracting firms, the results could be biased because contractor status is not exogenous: firms with the lowest cost of meeting affirmative actions requirements may be more likely to be contractors. Hence, we are left with an incomplete picture of both the impact of a controversial program and the potential consequences of its removal. What is needed, essentially, is a control group to which we can compare changes in outcomes for those affected by affirmative action.

The enactment of California Proposition 209 provides just such an opportunity. The measure, passed in the 1996 state elections and made effective in November of 1997, essentially outlawed existing local and state affirmative action programs in education, public hiring, and contracting, unless superseded by federal law. This change in state policy presents a natural experiment for measuring the labor market impact of removing of affirmative action programs. I use Current Population Survey (CPS) data to compare outcomes for minorities in California before and after affirmative action was removed to those same outcomes for white males. Then, to control for national trends in minority differentials, I compare this difference to the difference for a control group: states not undergoing similar changes in the law. The use of this triple difference technique to analyze the impact of removing affirmative action on employment, unemployment, labor force participation, and wages will provide evidence on the long-term effects of affirmative action.

2 The history and consequences of affirmative action policy

National legislation and impacts

Whereas equal employment opportunity (EEO) laws such as Title VII of the Civil Rights Act prohibit discrimination, affirmative action legislation goes further by requiring that proactive steps be undertaken to remedy inequalities produced by past discrimination. In 1965, President Johnson issued Executive

Order 11246, the primary regulation governing affirmative action, which requires that federal contractors "take *affirmative action* to ensure that applicants are employed, and that employees are treated during employment, without regard to their race, color, religion, sex, or national origin." Under its provisions, federal contractors must provide written affirmative action plans, progress reports, and submit to government compliance reviews. While EO 11246 only directly affected federal contractors, many state and local agencies and non-contractor private businesses voluntarily adopted similar programs in an attempt to address discrimination and avoid litigation under equal employment laws (Thomas and Garrett, 1999).

Early studies tend to indicate that affirmative action had a positive impact on the employment and occupational advancement of racial minorities.[1] Because of the inherent difficulty in separating gains from affirmative action from general trends in racial inequality, these studies rely on data from the Equal Employment Opportunity Commission (EEOC) to compare outcomes for firms that are federal contractors, and hence subject to federal affirmative action programs, with firms that are not. Ashenfelter and Heckman (1976) find that the demand for black males increased 3.3 percent more among contractors than non-contractors between 1966 and 1970. While they find similar employment gains for black males during the early seventies, Heckman and Wolpin (1976) and Goldstein and Smith (1976) find no improvement or even declines in employment for females at contractor establishments. However, as Leonard (1989) points out, affirmative action for women did not become stringently enforced until after the Equal Employment Act of 1972.

Studies of affirmative action in the late seventies and beyond tended again to find positive employment gains for racial minorities and additional, although smaller, gains for white females. Leonard (1984c) finds that between 1974 and

[1]For a survey of the literature on affirmative action, see Holzer and Neumark (2000a).

1980, contractor demand for black males grew 3.8 percent faster, demand for other minority males grew 7.9 percent faster, and the demand for white females grew 2.8 percent faster than that of non-contractors. Leonard (1984b) also finds that affirmative action appeared to have a relatively greater impact on minorities in skilled occupational groups, although Smith and Welch (1984) suggest that observed gains in occupational status may be due to contractors re-classifying jobs rather than to any real upward mobility. Rodgers and Spriggs (1996) find that the positive impact on employment continued through to 1992 for all groups except Hispanics, for whom they find a negative impact. Holzer and Neumark (2000b) have one of the few empirical studies with wide scope that does not depend on EEOC data. Using information from a survey of employers in four U.S. cities, they find that firms that use affirmative action do tend to recruit and hire more minorities and women. In fact, contrary to most earlier results, the use of affirmative action in hiring seems to have the largest effect for white women. For firms that report using affirmative action in hiring, the last employee hired is 8 percent more likely to be a white woman and 3 percent more likely to be a black man.

Benefits through employment gains and occupational advance, however, may mask underlying losses in efficiency. While the effects of affirmative action on market efficiency are not fully understood,[2] what evidence is present does not seem to suggest large declines in productivity. Leonard (1984a) combines EEOC data with industry level data and finds no evidence of lower productivity among federal contractors. In their study, Holzer and Neumark (2000b) find that while minorities and women hired under affirmative action appear to have lower readily observable qualifications, their employers do not report significantly lower performance for these groups than for white males. The authors suggest that this is the result of more intensive screening and training programs.

[2] Holzer and Neumark (2000a) suggest that this is an important area for future research

California legislation and impacts

While empirical studies have tended to focus on national legislation, state governments have also instituted equal employment laws and affirmative action programs. In 1959, five years before the passage of the federal Civil Rights Act, California passed the Fair Employment Practices Act, which outlawed discrimination in that state and created the Fair Employment Practices Commission (FEPC) (later given responsibility for housing as well and re-named the Fair Employment and Housing Commission) to enforce the act. The FEPC was also granted the power to "engage in affirmative action with owners" in order to remedy discrimination (State of California). In practice, the FEPC has been responsible for oversight of affirmative action plans for state contracts over $200,000. In addition, in 1974 California began requiring all public agencies to submit affirmative action reports to the State Personnel Board (SPB), which was responsible for the oversight and development of public affirmative action programs(Thomas and Garrett, 1999). In 1989 California established contracting set asides for minority and women-owned business, requiring that at least 15 percent of the total value of state contracts go to minority-owned businesses and 5 percent to women- owned businesses. So, prior to 1997, not only were federal employers and contractors in California subject to mandated affirmative action programs, but so were all public employees and state contractors.

However, attacks on these state programs in the mid 1990's have resulted in their formal dismantling. In 1995, then-governor Pete Wilson signed Executive Order 124-95, which directed state agencies to eliminate preferential treatments that exceed federal or state statutory requirements. Legally this could only apply to pre-standing executive orders, and thus should not have affected state affirmative action laws, but it is not clear, in practice, what effect it would have (Thomas and Garrett, 1999). A year later, California voters passed Proposition 209 outlawing all state affirmative action programs and hence releasing public employers as well as state contractors from affirmative action requirements. After lengthy court challenges, the new law went into effect in November, 1997.

While there has been a flurry of research on the impacts of Propositon 209 on higher education in California, economists have neglected to pay attention to the corresponding impacts on labor markets. Yet, given that 8 percent of California's work force is in the non-federal public sector[3] and nearly 15 percent of California small businesses claim California state and local governments as clients (Williams, 1999), we might expect Proposition 209 to affect more than educational institutions. On the other hand, Holzer and Neumark (1999) suggest that approximately 60 percent of firms are federal contractors and subject to federal affirmative action policy. So, while Proposition 209 is likely to have had an effect on public employers in California, it may have been considerably less binding on private firms that are still subject to federal law.

The Proposition 209 experiment

Not only does it seem reasonable to expect that this change in policy would have an impact on California labor markets, but it also provides an opportunity to address two shortcomings of the empirical evidence to date.

First, previous work has had to rely on the comparison of firms that participate in affirmative action to those that do not. Researchers have either used EEOC data to compare federal contractors to non-contractors or firm-level data to compare firms that report using affirmative action to those that do not. Yet, because firms self-select into using affirmative action (by choosing to be federal contractors or by voluntarily implementing their own programs), estimates of the impact of affirmative action may be biased downward. Federal contractors and voluntary participants may self-select precisely because it is relatively cheap to implement affirmative action. Moreover, the results of these studies

[3] This average is from the employment data used in this paper.

have only provided an indication of the firm or sector-level impact of affirmative action, not of its economy-wide impacts. For instance, it is known that minority employment was rising at both contracting and non-contracting firms that file EEO-1 reports (albeit more rapidly at the contracting firms), but what was happening at firms that do not have to provide data on their composition? Did this rise in employment mask a re-shuffling of minorities between sectors?

Second, there has been no previous opportunity to gauge the impact of removing affirmative action— only of implementing it. While we do not suffer from a shortage of theoretical models of affirmative action, there is comparatively scant evidence on its long-term consequences. Theoretically, any model of a binding and effective affirmative action program will predict that minority employment should rise while the policy is in place, leaving only the need to see empirically whether existing programs appear to be effective and what the extent of their impact is.

Depending on the assumptions made about the source of pre-existing inequality, affirmative action may or may not engender a long-term change in labor market differentials that would remain even if the program were removed. If labor market discrimination did not exist in the first place or if, as some models (e.g., Johnson and Welch, 1976) suggest, affirmative action is not an efficient policy, then removing affirmative action may cause the labor market to revert to its competitive equilibrium. On the other hand, certain models of discrimination do suggest a long-term impact for affirmative action. If, for example, labor market inequalities are the result of classic employer discrimination, then it is possible that by being forced to interact with minority groups, employer prejudices will diminish so that once affirmative action is removed there is no longer inequality. Alternatively, Coate and Loury (1993) consider a form of statistical discrimination in which employers are less likely to place minority workers in high skilled jobs because of negative stereotypes. As a result, minorities have less incentive to invest in human capital, leading to a self-fulfilling prophecy. Assuming that minority workers have the same fundamental ability, affirmative action could break this cycle and potentially create permanent change in negative stereotypes. A third theoretical alternative for predicting the continued effectiveness of affirmative action after its removal is that presented by Athey et al. (2000). In their model, entry level employees receive more mentoring from senior employees with similar characteristics. As a result, there is bias towards one type of employee in promotion that can be permanently broken by a temporary affirmative action program that introduces diversity.

The passage of Proposition 209 provides a natural experiment that can be used to address both shortcomings of previous studies. First, it provides a (presumably) exogenous shock to affirmative action policy that affects only workers in California, leaving workers in the rest of the country as a control group. Second, this is the first legislation that has attempted to dismantle affirmative action.[4] By comparing the relative change in labor market outcomes in California to the rest of the country, we can see what impact removing state-sponsored affirmative action had on women and minorities in California. If there was no impact, it could be the case that affirmative action was either ineffective in the first place in California or that it was effective in engendering long-term changes that remained even after its removal. If there was a negative impact on the employment of minorities, this suggests that either the prejudicial attitudes of employers were not changed under California's affirmative action program or that the program itself had engendered inefficiencies and reverse discrimination against whites.

[4]Other states and political entities followed suit after the proposal of Proposition 209. Washington state passed its own repeal of affirmative action in 1998 although similar proposals have failed elsewhere.

3 Data

I employ data from the outgoing rotation groups in the monthly Current Populaton Survey (CPS) from 1994-2001, placing emphasis on 1995, the year before the proposal of Proposition 209, and 1999, two years after the new law had gone into effect. Observations are dropped if an individual is employed but reports no hours or pay, reports unknown sector of employment, or is self-employed.[5] Observations from Washington state were also dropped because that state passed legislation similar to Proposition 209 in 1998.

The triple difference estimates in this analysis will rely on three divisions of the data. First, the observations are categorized as before or after the enactment of proposition 209 (e.g. 1995 or 1999, 1995 or 2000, and so on depending on the years being used). Second, individuals are divided into eight mutually exclusive and collectively exhaustive categories: white males, white females, black males, black females, other males, other females, Hispanic males, and Hispanic females. And third, the country is divided into two groups: an experimental state (California) and and the remaining control states or "nation."

Table 1 reports sample sizes for each cell. Because of its population, the sample sizes within California for even this detailed breakdown of minority groups are still fairly large. However, California is not necessarily representative of the country as a whole. It is more minority heavy than the rest of the country and has slightly lower rates of employment, but a similar distribution of employment across sectors and industry. The fact that California is more diverse than the country as a whole means that extrapolations from its experience with affirmative action to general predictions should be made cautiously.

[5] Because men are more likely than women to be self-employed, omitting this group tends to increase the number of women in the sample relative to the number of men

Average changes in labor force status and wages

Turning to the effects of affirmative action, Table 2 explores the change in non-participation in the labor force for white females after Proposition 209 was enacted. In 1995, 46.0 percent of white women in California over age 16 were not in the labor force while 32.1 percent of white males were not participating. In 1999, after Proposition 209 had gone into effect, the percentage of white females who were not in the labor force had fallen to 44.5 percent, but the participation of men showed a similar change. Overall, there was no significant change in the participation of white women relative to that of white men in California. As a control, I look at the same outcomes for the rest of the nation. Over the same period nationwide, the non-participation of white women had fallen by 1.3 percentage points relative to white men. Differencing these effects, relative to the rest of the country, non-participation among white females in California rose by 1.6 more percentage points than that for white males. However, this estimate is not significant.

In addition to women, racial minority groups in California may have also been affected by Proposition 209. Table 3 presents triple difference average changes for individual minority groups as well as for all minority groups together. The outcomes examined are the three labor force categories into which each individual falls: employed, unemployed, and not in the labor force as well as the hourly wage. For each group, the triple difference is calculated as in the preceding example. Note in particular that because employment, unemployment, and non-participation are mutually exclusive and collectively exhaustive, the relative changes for each group across these categories sum to 0. The point estimates indicate that, relative to white males and the rest of the nation, the proportion of minorities who were employed fell in California, but that there was little change in unemployment, leaving a rise in non-participation to account for most of the fall in employment. However, only the estimates for other males are significant. While there is also no significant change in wages, the point estimates are positive for six of the seven groups suggesting, for example, that the relative wage of white females rose by 76 cents. However, there is no clear prediction about the wage changes that might accompany a policy that can directly affect both wages and employment. It may be the case that the employees who are left are relatively more skilled, so average wages might rise. Alternatively, it might be the case that affirmative action also served to augment wage equality and so its removal might create a drop in wages. Given the possibility of opposing effects, if is not surprising to find no significant impact on wages.

Individual-specific differences

Although it is primarily viewed as a cross-sectional data set, the CPS can also be used as a panel in which each individual in the outgoing rotation group is observed twice. To construct a panel, I match the respondents at a particular address across year and then assume that the respondent is the same person if sex and race have not changed and if age has increased by 0 to 2 years. This allows approximately two thirds of the individuals in the outgoing rotation group in any given year to be matched to the previous year. However, because of a change in CPS methodology, matching is not possible for June-December of 1994 and 1995 and January-August of 1995 and 1996.

If Proposition 209 had an effect on the labor force status of women and minorities, then one would expect to find differences in the status of individuals across years. Moreover, by examining the change in outcome for the same individual, individual specific fixed effects (such as ability or skill) are eliminated. I examine the probability that an individual left the labor force between $t = 1$ and $t = 2$ given an observed change in labor force status but constructing an indicator variable that is 1 if he left the labor force and 0 if he entered the labor force. Conditioning on a change in labor force status reduces the sample size, but creates a binary variable for the first difference, assisting with inference for the double and triple differences.[6]

Table 4 reports the triple difference estimates of the relative probability that members of each minority group left the labor force given a change in participation. The first difference is the proportion of each group that left the labor force conditional on a change in participation. The second difference gives the proportion of each minority group that left the labor force relative to the proportion of white males. The third difference compares this change in California to the change in the rest of the nation. The estimates indicate that significant changes took place between 1994 and 1995 and 1995 and 1996.[7] Between 1994 and 1995 in California, white women were 16.7 percent more likely to have left the labor force given a change in participation than were white men relative to the nation as a whole. Black males, black females, and Hispanic females were also more likely to leave the labor force than to enter to it. Between 1995 and 1996, black women and Hispanic women were again more likely to leave the labor force than to enter it although the reverse is true for other females. As a whole, the estimates suggest a significant climb in the proportion of minorities who were leaving the labor force relative to entering it in the mid-nineties.

Note that the years for which these changes are observed are directly preceding or during the period when Proposition 209 was debated and passed. This could indicate an anticipation of the change in affirmative action policy,

[6]Consider, for instance, the difference for an individual in non-participation. It could be 1 (left labor force), 0 (no change), or -1 (entered labor force) and so is not binomially or normally distributed. Because of small sample sizes, it does not seem reasonable to invoke the Central Limit Theorem and non-parametric tests of differences for matched pairs are not appropriate for double or triple differences. By looking at whether an individual entered or left the labor force conditional on a change in participation, I create a binomial random variable and avoid these issues.

[7]These are the same years for which limited matching was possible due to change in CPS methodology. The significant estimates are noteworthy given the small sample sizes.

which seems plausible given the political environment in California at the time. While these results are indicative of significant change, they cannot be compared directly to estimates in upcoming sections which are based on cross sectional cuts of the data. In particular, the outcome variable in the regressions will not be conditioned on a change in participation and the span of time examined is longer than one year. Later results will provide similar evidence of a move out of the labor force, but this does not seem to be as concentrated in 1994-1995 as suggested here. However, these findings do bolster later results suggesting a significant change in participation.

4 Econometric model

To control further for the characteristics of the potential or actual labor force in estimating the impacts of Proposition 209, I turn to a triple difference regression framework. For simplicity, consider the case of only one minority treatment group. In the case of general outcome y, consider the equation

$$\begin{aligned} y_{ijt} = \mathbf{x}_{ijt}\boldsymbol{\gamma} + \beta_1 year_t + \beta_2 exper_j + \beta_3 treat_i + \beta_4 (year_t * exper_j) + \\ \beta_5 (year_t * treat_i) + \beta_6 (exper_j * treat_i) + \beta_7 (year_t * exper_j * treat_i) + \epsilon_{ijt} \end{aligned} \tag{1}$$

where $\mathbf{x}_{ijt}$ is a vector containing a constant and explanatory variables other than those that are part of the differencing, i indexes an individual, j indexes location, and t indexes time. In this case, $year$ is a dummy for the latter year in the regression (e.g. 1999 if we are comparing 1999 to 1995), $exper$ is a dummy indicating that the individual resides in California, the experimental state, and $treat$ indicates that the individual is a member of the minority treatment group. As is standard, the coefficient β_7 represents the triple difference estimate of the impact of Proposition 209 on outcome y for the treatment group.

A quick examination of the various differences of interest illustrates this. Note that the double difference estimate of the change in outcome for the treatment group in California relative to white males (wm) can be calculated as follows:

$$\begin{aligned}
\Delta_{treat,exper} &= y_{treat,exper,99} - y_{treat,exper,95} = \beta_1 + \beta_4 + \beta_5 + \beta_7 && (2)\\
\Delta_{wm,exper} &= y_{wm,exper,99} - y_{wm,exper,95} = \beta_1 + \beta_4 && (3)\\
\Delta^2_{exper} &= \Delta_{treat,exper} - \Delta_{wm,exper} = \beta_5 + \beta_7. && (4)
\end{aligned}$$

Similarly, the double difference for the control states is

$$\Delta^2_{nation} = \Delta_{treat,nation} - \Delta_{wm,nation} = \beta_5. \qquad (5)$$

And, so, the triple difference estimate of the impact of Proposition 209 is

$$\Delta^3 = \Delta^2_{exper} - \Delta^2_{nation} = \beta_7. \qquad (6)$$

I estimate a log wage regression to gauge the impact of removing affirmative action on hourly wages. Because no likely instrument is present for estimating a two-stage Heckman-type procedure, this is simply a wage regression *conditional* on employment. The possible biases that this may present are discussed along with the results in the following section.

The remaining three outcomes of interest– employment, unemployment, and non-participation– are binary variables and are commonly estimated with probit or logit models. As Ai and Norton (2003) point out, the marginal effect of the interacted variables in a nonlinear model is not the same thing as the marginal effect of the interaction term.[8] In other words, simply calculating the marginal effect of the triple difference term ignores the fact that one can-

[8]Nevertheless, the authors note that most studies with interaction terms in a nonlinear model report the marginal effect of the interaction term even though there is little intuitive explanation for what this means. In the differences-in-differences literature, Gruber (1994), for example, reports the marginal effect of the difference-in-difference coefficient.

not simultaneously "turn off" β_7 without affecting the other related interaction variables.

Some authors (e.g., Borjas, 2003) choose to use a linear probability model to avoid the complications that arise from a nonlinear model with interaction terms. However, linear probability models have their own (well-known) shortcomings. In this paper, I use a probit model and estimate the triple difference marginal effects through repeated differencing of the normal CDF. The double differences for the sample of individuals in the treatment group are:

$$\Delta^2_{i,exper} = [\Phi(\mathbf{x}_i\boldsymbol{\gamma} + \beta_1 + \beta_2 + \beta_3 + \beta_4 + \beta_5 + \beta_6 + \beta_7) - \Phi(\mathbf{x}_i\boldsymbol{\gamma} + \beta_2 + \beta_3 + \beta_6)] - [\Phi(\mathbf{x}_i\boldsymbol{\gamma} + \beta_1 + \beta_2 + \beta_4) - \Phi(\mathbf{x}_i\boldsymbol{\gamma} + \beta_2)] \quad (7)$$

and

$$\Delta^2_{i,nation} = [\Phi(\mathbf{x}_i\boldsymbol{\gamma} + \beta_1 + \beta_3 + \beta_5) - \Phi(\mathbf{x}_i\boldsymbol{\gamma} + \beta_3)] - [\Phi(\bar{\mathbf{x}}\boldsymbol{\gamma} + \beta_1) - \Phi(\bar{\mathbf{x}}\boldsymbol{\gamma})], \quad (8)$$

and the triple difference is

$$\Delta^3_i = \Delta^2_{i,exper} - \Delta^2_{i,nation}. \quad (9)$$

I average the triple difference marginal effects across individuals in the treatment group to get the average marginal effect.

The delta method is used to compute standard errors. Again, in previous work, authors have generally reported only the standard error for β_7, the triple difference probit coefficient. But, simply because the estimated probit coefficient on the triple difference term is significant does not mean that the marginal effect is. Let $h(\hat{\beta}, z_i)$ be the nonlinear function composed of the eight-fold differences of normal CDFs as expressed in Equation 10 and let V_β be the variance-covariance matrix of the coefficients. Then the variance for the average triple difference marginal effect is

$$V = E(\nabla_\beta h(\beta, x_i))V_\beta E(\nabla_\beta h(\beta, x_i))\prime. \tag{10}$$

This can be estimated by

$$\hat{V} = \left(\frac{1}{n}\sum_i \nabla_\beta h(\hat{\beta}, x_i)\right)\hat{V}_\beta\left(\frac{1}{n}\sum_i \nabla_\beta h(\hat{\beta}, x_i)\right)'. \tag{11}$$

Note that this method calculates the standard errors for the average triple difference marginal effect rather than the other commonly used option in evaluating marginal effects: the marginal effect for the average member of a group.[9]

5 Empirical analysis of Proposition 209

Labor force status

Table 5 reports the triple difference estimates of the average marginal effects of Proposition 209 on employment, unemployment, and non-participation.[10] In reality, the changes in employment, unemployment, and non-participation must sum to 0, but that added restriction is not placed on the marginal effects reported here. However, for the most part the sum of the unrestricted effects is fairly close to zero. In an attempt to identify possible short and longer term effects of the legislation, three pairs of years are examined: 1995 and 1999, 1995 and 2000, and 1995 and 2001.[11] Moreover, the results are presented for each of

[9] Ai and Norton (2003) provide an estimator for the variance of the marginal effect for the average individual when using interaction terms, but not for the average marginal effect. Thanks to Stephen Donald for his help with obtaining the correct estimator for the case used here.

[10] In addition to race and sex, age, marital status, interview month, education, region, urban status, citizenship, and nativity were also controlled for in all regressions. Wage regressions also included indicators of sector of employment (public, private, or federal), occupation, and industry.

[11] In all cases, 1995 is used as the base year to which post-legislation years are compared. The results are similar if 1993, 1994, or 1996 is used as the base instead. In addition, 1998 and 2002-2003 were also examined as post-legislation years. The triple difference coefficients are smaller in 1998, similar in 2002, and become insignificant in 2003. However, it is not clear how to interpret this since extending the time frame also increases the chance of unobserved events biasing the results.

the seven treatment groups as well as for all of the minorities together.

Looking at the results presented in Table 5, between 1995 and 1999 the relative employment of minorities fell by 2.8 percentage points while non-participation rose by 2.9 percentage points. Similarly, between 1995 and 2000 relative employment fell by 1.8 percentage points (but the change is not significant) and non-participation rose by 2.2 percentage points, and between 1995 and 2001 relative employment fell by 2.2 percentage points while non-participation rose by 2.0 percentage points. Breaking this down by group, between 1995 and 1999, relative non-participation rose by 2.9 percentage points for white females, 4.6 percentage points for black females, 1.4 percentage points for Hispanic males, 5.2 percentage points for Hispanic females, and 6.8 percentage points for other males. This increase in non-participation accounts for nearly all of the decline in employment for all groups except black females, who also saw a drop in unemployment. Only black males and other females do not exhibit significant changes in labor force status between 1995 and 1999. By 2001, however, there appears to have been a rise in non-participation for all minority groups except for black men, who show a significant fall in non-participation.[12]

As a whole, the results suggest that the impact of Proposition 209 was to move females and minorities from employment to out of the labor force. If, as the results indicate, the removal of affirmative action made it more difficult for women and minorities to find work, then this exit from the labor force is not surprising. Previous work has tended to indicate that women have more

[12]Interestingly, these estimates are nearly identical to those obtained using a linear probability model with robust standard errors, which avoids the complications inherent in estimating the marginal effects and their standard errors and makes it simple to restrict them to sum to zero. However, the estimates presented here are of much smaller magnitude than those obtained if the marginal effect of the interaction term is used instead. For instance, the average marginal effect of the triple difference coefficient– that is, the difference of the normal CDF with β_7 turned "on" and "off"– suggests that between 1995 and 1999, employment for all minorities fell by 10 percentage points, unemployment rose by 2.4 percentage points, and non-participation rose by 11.2 percentage points. That these effects are so much larger than those obtained using the correct method is somewhat alarming given the frequency with which marginal effects are calculated in this way.

elastic labor supplies than men and that they tend to be more responsive along the extensive participation margin (Blau and Kahn, 2005). In addition, when looking at the impact of minimum wage legislation, Mincer (1976) finds that affected groups tend to leave the labor force and, moreover, that females and minorities have relatively high participation elasticities.

Wages

As discussed previously, there is no clear prediction of the impact of removing affirmative action on wages. Relative wage changes will depend on the nature of pre-existing discrimination, the effectiveness of affirmative action, and the relative skill levels of the groups affected by its removal. It is thus not surprising that the results in Table 6 do not show such clear patterns as the labor force status results. No significant changes are observed between 1995 and 1999 with the exception of the relative wages of employed Hispanic females, which have risen by 5 percent. In 2000 we see a rise for other females, but the wages for black males have fallen by 9.2 percent. Between 1995 and 2001, only black females show a significant change in relative wages. As a whole, the results do not show a consistent effect for any of the groups. This could indicate that affirmative action had little effect on wages. Affirmative action laws, after all, did not directly address wage equality, which was covered by equal employment law. It could also be the result of skill selection among those leaving employment. Since the wage regressions are conditional on employment, the wages of those who remain employed could rise because they are relatively more skilled or fall because they are relatively less skilled than those who left.

Participation effects by age and education level

The wage findings do not provide consistent evidence of skill bias among those who remain employed, but they do not prove the contrary either. Previous studies have suggested that affirmative action helps to advance minorities into more skilled occupations (e.g., Goldstein and Smith, 1976; Leonard, 1984b). However, it is not clear how the removal of affirmative action would affect workers across skill groups. It could be the case that low-skilled workers become discouraged with the diminishment of opportunities for advancement and leave the labor force or, perhaps, it might directly impact those high-skilled workers who had previously advanced. In an attempt to gauge if a particular group is disproportionately affected by the legislation, the non-participation marginal effects were estimated for separate segments of the sample.

Table 7 reports these results. In columns (1)-(3), non-participation marginal effects are reported for three education levels: less than high school, a high school diploma, and education beyond high school.[13] On average, non-participation rose more for minorities with lower levels of education. For those with less than a high school degree, non-participation rose by 3.5 percentage points, for those with a high school degree or education beyond high school, the corresponding rises were 2.8 and 2.5 percentage points, respectively. This trend is evident for the disaggregated treatment groups as well, although, as in the earlier results, the rise tends only to be significant for white females, black females, hispanic females, and other males. Turning to age, columns (4)-(6) report results for three age brackets: 30 years old or younger, 30 to 50 years old, and older than 50. On average, non-participation rose more for the youngest group of workers than for older ones. When disaggregated by treatment group, for all groups except Hispanics it seems that non-participation rose more for individuals who are under 30 years old or over 50. This is not surprising given the expectation that very young and very old workers will be less attached to the labor force than middle aged workers. If education and age truly proxy for skill, then these findings suggest that the removal of affirmative action disproportionately

[13]These estimates are based on the NILF probit for the entire sample, but the marginal effects are calculated for the relevant level of education or age.

affected low-skilled workers.

Sector of employment

Previously I suggested that, because Proposition 209 does not supersede federal affirmative action laws, workers in California's public sector, who were covered by California policy but not federal, might see the largest effects from the measure. However, it is difficult to use CPS data to compare inter-sector differences. The results for the economy as a whole suggest that Proposition 209 did not affect the unemployment rate but did decrease participation. But, if an individual is not in the labor force, then we cannot identify what sector they may have worked in previously. I attempt to circumvent this problem by using a probit model to estimate the impact of Proposition 209 on the probability that an individual works in the private, public (state or local), or federal sector given that he is employed. Because some of the cell sample sizes become very small when disaggregated by both treatment group and sector, the estimation is performed only for all minorities as a single treatment group.

Table 8 presents these results. The point estimates suggest that minorities were actually slightly more likely to work in the public sector and less likely to work in the private sector after the removal of affirmative action. This is counter to the expectation that the negative effects of Proposition 209 would be strongest for state and local workers in California. It may be the case that private employers did respond to the removal of state-sponsored affirmative action in California and to the general anti-affirmative action climate of the period. While many private-sector firms in California were likely to be federal contractors, federal affirmative action policy had also been under intense legal scrutiny during the nineties and private employers in California may have felt more bound by state policy than federal. However, the coefficients here are insignificant and of small magnitude, thus providing no strong evidence of a significant shift between sectors of employment.

6 Conclusion

The enactment of Proposition 209 in California created a unique opportunity to study the labor market effects of the removal of affirmative action programs. Changes in minority outcomes in California relative to those of white males are compared to the same differences for the rest of the nation in order to separate the effects of Proposition 209 from general trends in inequality.

The results suggest that there was a sharp drop in employment after the passage of Proposition 209, which resulted in minorities leaving the labor force. Between 1995 and 1999, relative employment for minorities in California fell by 2.8 percentage points as these groups left the labor force, causing non-participation rates to climb by a corresponding 2.9 percentage points. Disaggregated, non-participation rates in California rose by 2.9 percentage points for white women, 4.6 percentage points for black women, 5.2 percentage points for Hispanic women, 1.4 percentage points for Hispanic men, and 6.8 percentage points for other men. There appears to have been little corresponding change in wages rates, but this may be driven by skill bias in the workers who remained employed in later years.

The decline in minority participation raises doubts about affirmative action programs. It is consistent with one of two hypotheses: that affirmative action is inefficient and creates reverse discrimination or that affirmative action is ineffective at engendering permanent change in prejudices that create labor market inequality. A final possibility is that California's affirmative action programs had not been in place long enough to engender permanent alteration in inequality. However, given that California had pursued affirmative action for over a generation, this may be equally discouraging.

Table 1: **Sample Sizes**

	1995		1999	
	Nation	California	Nation	California
white male	94,335	5,113	84,120	5,191
white female	112,672	6,045	98,649	5,927
black male	11,582	595	10,238	597
black female	16,689	773	14,409	812
hispanic male	7,301	2,805	8,315	2,992
hispanic female	8,533	3,144	8,991	3,168
other male	4,857	1,310	4,431	1,246
other female	5,758	1,500	5,198	1,486
total	261,727	21,285	234,351	21,419

Table 2: **Change in Non-Participation, 1995-1999: White Males and Females**

	1995	1999	Time Difference for Group
California			
white females	0.460	0.445	-0.015
	(0.006)	*(0.006)*	*(0.009)*
white males	0.321	0.302	**-0.018**
	(0.007)	*(0.006)*	*(0.009)*
Group difference for a given year	**0.140**	**0.143**	
	(0.009)	*(0.009)*	
Double Difference	0.003		
	(.013)		
Nation			
white females	0.432	0.421	**-0.012**
	(0.001)	*(0.002)*	*(0.002)*
white males	0.284	0.285	0.001
	(0.001)	*(0.002)*	*(0.002)*
Group difference for a given year	**0.148**	**0.135**	
	(0.002)	*(0.002)*	
Double Difference	**-0.013**		
	(0.003)		
Triple Difference	0.016		
	(0.013)		

*Standard errors are in parentheses below estimates. All differences in bold are significant at the 5% level. Rounding is done after calculations.

Table 3: **Triple Differences Summary, 1995-1999**

	Employed	Unemployed	Not in Labor Force	Hourly Wage
white males	-	-	-	-
white females	-0.021	0.005	0.016	0.761
black males	-0.054	0.008	0.047	-0.240
black females	-0.034	-0.019	0.053	0.256
hispanic males	-0.002	-0.010	0.012	0.185
hispanic females	-0.029	-0.004	0.033	0.504
other males	**-0.060**	0.003	**0.057**	0.028
other females	-0.016	0.002	0.014	0.765
all minorities	-0.017	-0.003	0.020	0.012

*Values in bold are significant at the 5% level. All monetary values are in 1995 dollars.

Table 4: **Triple differences for proportion of each group that left the labor force conditional on a change in participation**

	93-94	94-95	95-96	96-97	97-98	98-99
white males	-	-	-	-	-	-
white females	0.015	**0.167**	0.155	0.017	-0.007	0.038
black males	-0.039	**0.477**	0.251	0.010	0.019	-0.013
black females	-0.156	**0.289**	**0.318**	-0.061	0.052	0.075
hispanic males	-0.017	0.177	-0.072	-0.018	-0.041	-0.033
hispanic females	-0.036	**0.186**	**0.235**	-0.009	-0.010	-0.040
other males	0.015	0.068	0.100	0.046	-0.127	0.006
other females	-0.054	0.155	**-0.276**	0.048	-0.022	-0.032
all minorities	-0.038	**0.157**	0.116	-0.008	-0.034	-0.022

*Values in bold are significant at the 10% level.

Table 5: **Triple Difference Marginal Effects for Employment, Unemployment, and Non-Participation Probits**

	1995-1999		1995-2000		1995-2001	
	coef	*se*	*coef*	*se*	*coef*	*se*
Employment						
white females	**-0.034**	0.011	-0.018	0.011	**-0.026**	0.011
black males	-0.035	0.027	-0.012	0.027	0.027	0.027
black females	-0.029	0.024	**-0.054**	0.024	**-0.051**	0.024
hispanic males	-0.008	0.016	**-0.030**	0.016	**-0.034**	0.016
hispanic females	**-0.048**	0.017	**-0.045**	0.017	**-0.034**	0.017
other males	**-0.071**	0.020	**-0.047**	0.020	**-0.068**	0.001
other females	-0.020	0.020	-0.022	0.020	-0.028	0.020
all minorities	**-0.028**	0.010	-0.018	0.010	**-0.022**	0.010
Unemployment						
white females	0.004	0.004	-0.001	0.004	0.003	0.004
black males	0.009	0.014	0.011	0.014	0.002	0.014
black females	-0.016	0.011	**-0.022**	0.010	0.002	0.012
hispanic males	-0.003	0.008	0.004	0.008	0.006	0.008
hispanic females	-0.003	0.007	0.005	0.007	$<$0.001	0.008
other males	0.005	0.009	-0.009	0.008	0.010	0.009
other females	$<$0.001	0.007	$<$0.001	0.006	0.008	0.007
all minorities	$<$0.001	0.004	-0.003	0.004	0.003	0.005
Non-participation						
white females	**0.029**	0.010	0.018	0.010	**0.022**	0.011
black males	0.029	0.026	0.002	0.026	**-0.024**	0.026
black females	**0.046**	0.023	**0.079**	0.024	**0.050**	0.024
hispanic males	**0.014**	0.015	0.029	0.015	**0.032**	0.015
hispanic females	**0.052**	0.017	**0.041**	0.016	**0.036**	0.016
other males	**0.068**	0.020	**0.059**	0.019	**0.059**	0.019
other females	0.020	0.020	0.021	0.019	**0.019**	0.019
all minorities	**0.029**	0.009	**0.022**	0.009	**0.020**	0.010

*Marginal effects are averages across the relevant treatment group. Marginal effects and standard errors are calculated as outlined in the text. Values in bold are significant at the 5% level.

Table 6: **Triple Difference Coefficients for Log(Wage) Regressions**

	1995-1999		**1995-2000**		**1995-2001**	
	coef	*se*	*coef*	*se*	*coef*	*se*
white females	0.001	0.018	-0.014	0.019	-0.003	0.020
black males	-0.039	0.039	**-0.092**	0.040	-0.005	0.040
black females	-0.029	0.040	-0.029	0.037	**-0.100**	0.041
hispanic males	0.007	0.021	0.005	0.021	-0.019	0.021
hispanic females	**0.052**	0.024	0.044	0.023	0.037	0.025
other males	-0.016	0.033	0.044	0.031	0.009	0.033
other females	0.041	0.030	**0.087**	0.029	0.019	0.032
all minorities	0.010	0.015	0.008	0.015	-0.002	0.016

*Standard errors are robust. Values in bold are significant at the 5% level. All monetary values are 1995 dollars.

Table 7: **Triple Difference Marginal Effects by Education and Age for Non-Participation Probit, 1995-1999**

	Education			**Age**		
	< High School	High School	> High School	<30	30-50	> 50
white females	**0.032**	**0.030**	**0.027**	**0.036**	**0.026**	**0.028**
black males	0.032	0.031	0.027	0.035	0.026	0.030
black females	**0.049**	0.048	0.045	**0.056**	0.041	**0.046**
hispanic males	0.016	0.014	0.012	0.016	0.014	0.010
hispanic females	**0.057**	**0.050**	**0.045**	**0.058**	**0.052**	**0.043**
other males	**0.083**	**0.072**	**0.062**	**0.080**	**0.060**	**0.067**
other females	0.025	0.021	0.017	0.023	0.016	0.022
all minorities	**0.035**	**0.028**	**0.025**	**0.035**	**0.028**	**0.025**

*Marginal effects are averages across the relevant treatment group and are calculated as outlined in the text. Values in bold are significant at the 5% level.

Table 8: **Triple Difference Marginal Effects for Sector of Employment Probits for all Minorities, 1995-1999**

	coef	*se*	*p-value*
P(Private|Emp)	-0.0011	0.0006	0.0805
P(Public |Emp)	0.0012	0.0008	0.1423
P(Federal|Emp)	-0.0001	<0.0001	0.1206

*Marginal effects are averages across the relevant treatment group and are calculated as outlined in the text. Values in bold are significant at the 5% level.

References

Chunrong Ai and Edward C. Norton. Interaction terms in logit and probit models. *Economics Letters*, 80, July 2003.

Orley Ashenfelter and James Heckman. Measuring the effect of an antidiscrimination program. In Orley Ashenfelter and James Blum, editors, *Evaluating the Labor-Market Effects of Social Programs*. Industrial Relations Section, Princeton University, 1976.

Susan Athey, Christopher Avery, and Peter Zemsky. Mentoring and diversity. *American Economic Review*, 90, September 2000.

Francine D. Blau and Lawrence M. Kahn. Changes in the labor supply behavior of married women: 1989-2000. NBER working paper 11230, 2005.

George J. Borjas. Welfare reform, labor supply, and health insurance in the immigrant population. *Journal of Health Economics*, 22, November 2003.

Stephen Coate and Glenn C. Loury. Will affirmative action policies eliminate negative stereotypes? *American Economic Review*, 83, December 1993.

Morris Goldstein and Robert S. Smith. The estimated impacts of the antidiscrimination program aimed at federal contractors. *Industrial and Labor Relations Review*, 29, July 1976.

Jonathan Gruber. The incidence of mandated maternity benefits. *The American Economic Review*, 84, June 1994.

James J. Heckman and Kenneth I. Wolpin. Does the contract compliance program work? An analysis of Chicago data. *Industrial and Labor Relations Review*, 29, July 1976.

Harry Holzer and David Neumark. Are affirmative action hires less qualified? evidence from employer-employee data on new hires. *Journal of Labor Economics*, 17, July 1999.

Harry Holzer and David Neumark. Assessing affirmative action. *Journal of Economic Literature*, 38, September 2000a.

Harry Holzer and David Neumark. What does affirmative action do? *Industrial and Labor Relations Review*, 53, January 2000b.

George E. Johnson and Finis Welch. The labor market implications of an economywide affirmative action program. *Industrial and Labor Relations Review*, 29, July 1976.

Jonathan S. Leonard. Antidiscrimination or reverse discrimination: The impact of changing demographics, Title VII, and affirmative action on productivity. *The Journal of Human Resources*, 19, Spring 1984a.

Jonathan S. Leonard. Employment and occupational advance under affirmative action. *Review of Economics and Statistics*, 66, August 1984b.

Jonathan S. Leonard. The impact of affirmative action on employment. *Journal of Labor Economics*, 2, October 1984c.

Jonathan S. Leonard. Women and affirmative action. *Journal of Economic Perspectives*, 3, Winter 1989.

Jacob Mincer. Unemployment effects of minimum wages. *Journal of Political Economy*, 84, August 1976.

State of California. Fair employment and housing act. Title 2 Division 3 Part 2.8 of California government code, section 12988 (1959-2005).

Executive Order no. 11426 3 C.F.R. 339 (1964-1965).

William M. Rodgers, III and William E. Spriggs. The effect of federal contractor status on racial differentials in establishment-level employment shares: 1979-1992. *The American Economic Review*, 86, May 1996.

James P. Smith and Finis Welch. Affirmative action and labor markets. *Journal of Labor Economics*, 2(2), April 1984.

Ward Thomas and Mark Garrett. U.S. and California affirmative action policies, laws, and programs. In Paul Ong, editor, *Impacts of Affirmative Action: Policies and Consequences in California.* Alta Mira Press, 1999.

Darrell L. Williams. The California civil rights initiative: Which firms stand to lose and how much? In Paul Ong, editor, *Impacts of Affirmative Action: Policies and Consequences in California.* Alta Mira Press, 1999.