

## **Interracial Workplace Cooperation: Evidence from the NBA**

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### **Abstract**

Using data from the National Basketball Association (NBA), we examine whether patterns of workplace cooperation occur disproportionately among workers of the same race. We find that, holding constant the composition of teammates on the floor, basketball players are no more likely to complete an assist to a player of the same race than a player of a different race. Our confidence interval allows us to reject even small amounts of same-race bias in passing patterns. We find some evidence of own-race bias in situations where the outcome of a particular play is less important. Our findings suggest that high levels of interracial cooperation can occur in a setting where workers are operating in a highly visible setting with strong incentives to behave efficiently.

JEL Codes: J15, J71, L23

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## **Introduction**

Recent research shows that when individuals are forced to make quick decisions, they often exhibit same-race preferences, even if they are unwilling to admit to biased racial attitudes. For example, Price and Wolfers (2010) show that NBA referees are more likely to call fouls against players of a different race than players of their own race, and Parsons et al. (2011) find that umpires are more likely to call strikes for pitchers of their own race. Similarly, Antonovics and Knight (2005) find that police are less likely to search the vehicle of someone of their own race, and Donohue and Levitt (2001) find that an increase in the number of police of a certain race is associated with an increase in arrests of people of the other race.

This same-race bias could play an important role in collaboration among colleagues in a workplace. For example, managers might be more likely to give favorable assignments to same-race employees. Alternatively, colleagues may depend disproportionately on same-race colleagues for advice or help. Collectively, such decisions may reduce the workplace productivity and satisfaction of employees of a minority race. These decisions may play a role in explaining the extent of workplace segregation (Hellerstein and Neumark 2008). Furthermore, this bias would undermine the argument that productivity is higher in groups that are ethnically diverse (Page 2007).

In this paper, we examine the effects of group heterogeneity on teamwork by studying specific and measurable actions within teams. In traditional firm-level data, it is often difficult to obtain measures of cooperation. As a result, we use play-by-play data from the National Basketball Association (NBA). These data allow us to determine for each basket, who passed the ball and which other players were on the court at the time.

We develop a simple model which allows the optimal pass to depend on the particular combination of teammates on the floor. We then test whether the pattern of observed assists demonstrates evidence of same-race bias.

We find no evidence that, conditional on the set of teammates on the court, players are more likely to pass to a teammate of their same race. Our baseline empirical strategy controls non-parametrically for the joint distribution of shot quality for all teammates on the floor. In other words, we account for differences in ability across teammates. Furthermore, the shooting opportunities for one teammate are allowed to depend arbitrarily on the set of other teammates on the floor. Robustness checks, in which we reduce the flexibility of our empirical specification to increase statistical precision, yield the same substantive results. Our evidence suggests that cross-race cooperation may not be a problem in workplaces where employees have common goals and extensive experience working with each other.

### **Similarity and Cooperation**

There is considerable research, empirical and theoretical, indicating that diversity can lead to improved economic outcomes (Alesina and La Ferrara 2005; Hong and Page 2004; Alesina, Spolaore, and Wacziarg 2000). These gains from diversity depend on the various groups being willing to cooperate. This may explain why other studies have found that increased racial diversity is associated with lower group performance (Kurtulus 2008; Timmerman 2000). In this paper, we expand the literature on the effects of group heterogeneity on outcomes by studying specific and measurable actions within teams.

Our analysis of cooperation is one form of own-race bias that has been documented in other types of interaction including referee-player (Price and Wolfers 2010; Parsons et al. 2011), employer-employee (Stoll, Raphael, and Holzer 2004) and officer-offender (Antonovics and Knight 2004, Donohue and Levitt 2001). What distinguishes cooperation from these other settings is that players are working together towards a common objective, while these other settings involve a more adversarial relationship. In addition, there is research showing that the racial composition of one's group affects decisions similar to cooperation, such as willingness to provide a public good (Martinez-Vazquez, Rider, and Walker 1997), form a coalition (Brasington 1999), or increase welfare spending (Luttmer 2001).

### **NBA Data**

Our analysis draws on play-by-play data that we collected from espn.com for all regular season and playoff games during October 2002-June 2008. The data includes an entry for every occurrence during the game that might be important for compiling game-level statistics. For each shot that is completed on the court, the data provide the name of the person who shot the basket and the person who was awarded with an assist, if the shot was assisted (58.9% are). Using the timing of substitutions in the data, we apply recursive methods to determine the ten players on court at any given time.

Thus, for each assisted shot we know who the passer was and who the other four available players would have been. For each of these players we include information about their race, position, and other characteristics. The player's race information comes from data collected by Kahn and Shah (2005), Price and Wolfers (2010), and our own

coding from more recent online photos of the players. Our racial coding is based on a simple measure of “black” or “not-black” (we refer to the not-black category as white and it includes non-black Asian and Hispanic players).

As part of our empirical strategy, we construct identifiers for each unique group of four players available to receive a pass on a particular play. For our primary analysis, we limit our sample to passing opportunities in which there was at least one player of each race available to receive the pass. This eliminates about 39.3% of our observations for situations in which the four players available to receive the pass are black and another 0.32% of our observations when those four players are white. Including these observations in our sample would bias our estimates of own-race discrimination towards zero, since the passer in these situations has no choice regarding the race of the player he can pass to<sup>1</sup>.

Our estimates using player-group fixed effects rely on variation in the race of the passer. As a result, we also eliminate each unique grouping of the four players in which all of the passers that we observe for that group are the same race. This restriction eliminates an additional 11.1% of our original sample. The combination of these two restrictions reduces our original sample of 302,714 assisted shots to 149,059 observations. These sample restrictions leave a final sample that includes more observations from teams with a higher fraction of white players.<sup>2</sup> However, we also run our results with the full sample of observations and find results that are very similar to

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<sup>1</sup> Since our results indicate that there is no racial bias, we want to use the empirical strategy that has the best chance of detecting racial bias if it were to exist.

<sup>2</sup> For example, only 20.3% of the observations for the Utah Jazz are excluded (a team where 56% of the shots are made by black players) while 81.2% of the observations for the Detroit Pistons are excluded (a team where 92.8% of the shots are made by black players).

those of our primary sample (with coefficients which are slightly smaller in most cases, as expected).

Table 1 shows summary statistics for both our restricted and unrestricted samples. The primary difference between the full and restricted sample is the fraction of shooters that are white. White players completed 37.1 percent of the shots and made 20.6 percent of the passes in our sample.<sup>3</sup> Separating the probability that the shooter is white based on the race of the passer (as done in columns 4 and 6) provides an initial estimate of racial gaps in passing. A white passer is 3.7 percentage points more likely to complete an assist to a white shooter than is a black passer. However, this simple estimate fails to capture any clumping of white players on the same team (or on the court together) or differences in the positions they play.

### **Model and Empirical Strategy**

Before progressing to our empirical specification and findings, it is helpful to outline a simple economic model of cooperation. Basketball involves complex offensive and defensive strategies. For this reason, we define a simpler game that will highlight the intuition involved and suggest an empirical strategy for identifying possible same-race bias in passing patterns.

Consider a game with five players. One of the five players is initially endowed with the ball. The player then passes the ball to one of the four remaining players with the best shot at the basket. We define  $S$  as the set of four players available for the pass.

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<sup>3</sup> These are an overestimate of the actual fraction of passes and shots made by white players because our sample is limited to observations in which there was at least one player of each race available to receive the pass. Without this restriction, white players would make 21.5% of the shots which is roughly in line with their representation in the NBA.

Player  $i \in S$  has an opportunity for shot, the quality of which is given by  $\mu_i$ . Player  $j \in S$  is another player in the passer's choice set with shot quality  $\mu_j$ . We do *not* assume that  $\mu_i$  and  $\mu_j$  are independent or identically distributed. We do, however, assume that both are independent of the player passing the ball. We assume that the player with the ball passes to his teammate with the highest quality shot.<sup>4</sup> Given these assumptions, the probability that the player with the ball passes to player  $i$  given the set of available passing options,  $S$ , can be written  $\theta_{i,S} = \text{prob}(\mu_i > \mu_j \forall j \neq i \in S)$ .<sup>5</sup>

This simple model suggests a tractable empirical specification to test the role of race in on-the-job cooperation. We can estimate the following linear probability model:

$$(1) \quad 1_{S,p}^{\text{scorer=white}} = \theta_{w,S} + \beta 1_{S,p}^{\text{assister=white}} + \varepsilon_p,$$

Where  $1_{S,p}^{\text{scorer=white}}$  is an indicator variable that takes on a value of 1 if the player scoring the basket is white, given the set of available scorers  $S$ , during possession  $p$ .  $\theta_{w,S}$  is estimated by a set of dummy variables for every combination of four players available to receive the assist. This controls non-parametrically for the probability that a white player has the best shot, taking into account the joint distribution of shot quality among all players eligible to receive a pass. Thus,  $\theta_{w,S}$  accounts both for the talent of every player available for a pass, as well as how the players interact while on the floor together. It

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<sup>4</sup> It could also be true that the player takes into account not only shot quality and racial preference but also the preference of the coach and the social pressure of the fans. In this case we would simply reinterpret shot quality as a measure of the total utility benefit associated with passing to a particular player.

<sup>5</sup> As an example of this, consider the case where  $\mu_i = \bar{\mu}_i + \varepsilon_i$ , where  $\bar{\mu}_i$  represents fixed player ability and  $\varepsilon_i$  is distributed according to a Type 1 extreme value distribution. In this case, the probability that the ball

is passed to player  $i$  is given by  $\text{prob}(\mu_i > \mu_j \forall j \neq i) = \frac{\exp(\bar{\mu}_i)}{\sum_{j=1}^4 \exp(\bar{\mu}_j)}$ .

also implicitly controls for the position of every potential pass recipient.  $1_{S,p}^{assister=white}$  takes on a value of 1 if the player making the assist is white, and  $\varepsilon_p$  is the residual.

Identification arises from the fact that both a white and a black player choose among the same set of teammates when passing the ball. If white and black passers are solving the same optimization problem with the same constraints,  $\beta$  should be statistically insignificant from zero. A non-zero coefficient suggests that the race of the passer and potential scorers affects the pattern of assists and hence on-the-job cooperation.

The disadvantage of our preferred approach is that it consumes literally tens of thousands of degrees of freedom since we include separate fixed effects for sets of players that differ only by a single role-player. In doing so, we discard large amounts of potentially useful information. To the extent that we can approximate  $\theta_{w,S}$  without the inclusion of so many dummy variables, we will increase the precision of the estimated same-race bias.

To do so, we calculate the fraction of assisted baskets scored by a player while on the floor, excluding baskets in which that player made the assist. This is calculated separately for each season. For player  $i$ , we denote this probability  $\hat{\pi}_i$ . This reflects the historical probability that player  $i$  has the best shot. We can sum these probabilities across all white players on the floor at a point in time to approximate the probability that a white player has the best shot at that particular point in time. Mathematically, this proxy for  $\theta_{w,S}$  is given by:

$$(2) \quad \hat{\theta}_{w,S}^{proxy1} = \sum_{i \in W \cap S} \hat{\pi}_i,$$



where  $W$  is the set of white players and  $S$  is the set of potential pass recipients on the court. While this measure is simple, it fails to take into account any interactions between players on the court. A symptom of this is that even if all potential pass recipients were white, the measure would almost certainly be above or below one.

For this reason, we construct a second proxy by normalizing this measure by the propensity of all players in the choice set to score off an assist. This measure again represents a probability that a white player has the best shot but has been normalized to lie between zero and one. This second proxy is given by:

$$(3) \quad \hat{\theta}_{w,S}^{proxy2} = \frac{\sum_{i \in W \cap S} \hat{\pi}_i}{\sum_{j \in S} \hat{\pi}_j}.$$

This proxy takes into account that a player's probability of scoring depends not only on his skill but the skill of the other teammates on the floor as well. It fails to take into account, however, all of the possible idiosyncratic interactions between teammates the way our preferred approach does.

Using the second of these two proxies for the probability that a white player has the best shot, we supplement our primary findings by estimating linear probability models of the following form<sup>6</sup>.

$$(4) \quad 1_{S,p}^{scorer=white} = \alpha_0 + \alpha_1 \hat{\theta}_{w,S}^{proxy2} + \beta 1_{S,p}^{assister=white} + \varepsilon_p$$

If our proxy performs well, we would expect our estimate of  $\alpha_1$  to be close to one.

Under the null hypothesis of no same-race bias, we would still expect  $\beta$  to be close to zero.<sup>7</sup>

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<sup>6</sup> We estimate all the regressions using logit (table 3) and conditional logit (table 2) regressions. The marginal effects and standard errors are nearly identical to those using linear probability models.

## Findings

As a starting point, we estimate a simple regression model in which we analyze the relationship between the passer's race and the race of the shooter. All of the regressions we describe in this section include controls for the passer's position and whether he is in his first year with the team. In this very simple regression, we find that the probability that the shooter is black increases by about 3.53 percentage points when the passer is white. This is similar to the raw difference that we observed in Table 1.

This simple regression fails to take into account the fact that there is racial clumping of players in the NBA with some teams having a lot of white players and others have few. When we control for the number of black players (besides the passer) on the court, we find that this estimated racial difference disappears completely (with an insignificant point estimate of 0.26 percentage points). We also get a very small estimate of racial bias when we include team-year fixed effects (-0.44 percentage points).<sup>8</sup>

Ideally, we want to control for the probability that a white player has the best shot, taking into account the joint distribution of shot quality among all players eligible for a pass. We do this by including a fixed effect for each four-player combination on the court aside from the passer. We find that the probability that the shooter is white increases by 0.65 percentage points relative to a baseline probability of 37.1 percentage points (or a 1.8% increase) and is not statistically significant. Even at the upper end of the

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<sup>7</sup> In all specifications, we cluster the correct standard errors at the team level. This takes into account that passing decisions may not be independent across assist opportunities. For example, a particular player may consistently look to pass to a certain teammate for reasons independent of the race of the two players.

<sup>8</sup> If we were to include the team-year fixed effects without controls for the number of black players on the court, we would get a large negative coefficient (-5.37 percentage points). This simply reflects the mechanical bias that, on a given team, a white passer will have fewer white players than black players to pass to. This is similar to the problem that Guryan, Kroft, and Notowidigdo (2009) discuss in estimating peer effects in golf.

95% confidence interval, a white passer is only 1.8 percentage points more likely to pass to a white teammate than a black passer.<sup>9</sup>

Our player-group fixed-effects model has the disadvantage of consuming thousands of degrees of freedom since we include separate fixed effects for sets of players that differ by only one person. In column 5, we provide the results of an alternative specification in which we approximate the probability that a white player has the best shot ( $\theta_{w,s}$ ) as shown in equation (3). As expected, we find that the coefficient on our proxy of  $\theta_{w,s}$  is relatively close to one, indicating that our constructed measure effectively captures the probability that a white player has the best shot (it is even closer to one when we use the full sample in Table 3).

Our test of own-race bias is based on whether deviations from this predicated probability are influenced by the race of the passer. We find no evidence of a preference by players to pass to players of their own race. Our estimate in the first column indicates that the probability that a white player receives an assist is only 0.04 percentage points higher when the passer is white than when the passer is black. In addition, the standard errors are about a third smaller than in our fixed-effects model, providing an even smaller upper bound estimate of the amount of own-race bias.

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<sup>9</sup> While we find no effect of same-race bias in passing patterns, an apparent effect would not necessarily indicate on-court animus. Instead, a high incidence of within race passing might indicate familiarity between players of the same race. In other words, a white player might be happy to make a pass to a black player in position to score but instead passes to a white teammate because he knows better where that teammate will be on the floor. In this case, race-based social investments off the court translate into low levels of cross-racial cooperation off the court. While the proximate cause of the racial bias differs, the ultimate consequence is the same.

## **Robustness Checks**

One concern with all specifications is that we have data on only completed assists. Levels of racial cooperation could also affect the number of unassisted baskets as well as passing patterns that do not lead to assists. Fortunately, we are able to test the hypothesis that probability of an unassisted basket depends on the racial composition of a player's teammates. In particular, we estimate a linear probability model in which the dependent variable is the probability that a completed shot was assisted.

In our preferred specifications, we control for the race and position of the scorer as well as the fraction of that player's baskets that were assisted over the course of the entire season. This takes into account that some players have higher unconditional probabilities of scoring unassisted baskets. We also include a fixed effect for the set of other players who are on the court at a given time. The variable of interest is the interaction between the white shooter variable and the number of black players on the court.

Under the racial animus hypothesis, the coefficient should be significantly positive. In all specifications, the coefficient is close to zero. Once we control for either team-year or teammate combination fixed effects, the coefficient is statistically insignificant. Overall, these specifications, which are available in the online appendix, are consistent with high levels of cross-racial cooperation.

We cannot, however, observe passes that were made that did not lead to shots and potential assists that were not converted. While this data limitation could affect the apparent magnitude of same-race bias, it does not affect the sign of the coefficient. Suppose a player systematically passes to teammates of his own race, even though they

have worse shots than teammates of another race. In this case, assists between players of the same race will be relatively more common than assists between players of differing races. This difference in assists will be less than the difference in attempted assists, however, because passes made for race-based reasons will be less likely to lead to converted baskets. When players choose to score alone or make a pass to a teammate out of position to score instead of assisting to a teammate of a differing race, this also increases the relative frequency of same-race assists. Thus data limitations may affect the magnitude of the observed same-race bias, but our procedure still provides a valid test for the existence of same-race passing preferences.

An additional possibility is that passers discriminate differentially against the worst players of the other race. If this is true, we would expect lower levels of interracial cooperation when the same race players are better shooters than other race players. To test this hypothesis, we re-estimate our preferred specification but include an interaction between our white passer indicator variable and the difference in field goal percentage between the potential white and black pass recipients. The coefficient on this interaction and the white passer variable are both insignificant, suggesting that racial differences in shooting percentage do not appear to affect the level of own race bias.

Recent research by Parsons et al. (2011) indicates that racial bias on the part of baseball umpires is price sensitive. The umpires exhibit more bias in situations where there is less scrutiny of their decision or when the outcome of their decision is less important. We examine whether own-race bias in passing is influenced by the importance of the situation by focusing on passes during the fourth quarter and splitting our sample based on the score margin at the start of the quarter.

Rather than pick an arbitrary cutoff, we find the estimated coefficient and confidence interval for each possible cutoff (i.e. games in which one team is up by at least X points at the start of the fourth quarter). The solid line in Figure 1 provides the estimated effect of own-race bias and the shaded area provides the 95% confidence interval. Figure 1A is based on estimates using the player-group fixed effects and Figure 1B is based on estimates using our theta-proxy model. While the precision of our estimates is much tighter when using the theta-proxy model, both figures provide suggestive evidence that the amount of own-race bias increases when the importance of making a particular pass is lower.<sup>10</sup> In fact, when using the theta-proxy approach, we find statistically significant levels of own-race bias when using a cutoff of a score margin between 11 and 17 points at start of the fourth quarter.

The prior results suggest that discrimination may be more likely when it is unlikely to affect the outcome of a game. It may also be possible that discrimination is practiced most by star players whose employment in the NBA is secure even if they exhibit discriminatory behavior that adversely affects performance.<sup>11,12</sup> Marginal players have little leeway to exhibit discriminatory behavior, since doing so may reduce their effectiveness and cost them their jobs. We test this hypothesis by running separate regressions for players who averaged more than 30 minutes per game and players who

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<sup>10</sup> To further isolate the effect of the game's importance on racial bias, we ran the same regressions as in column (3) of Tables 2 and 3, but split the sample based on whether or not the team performing the assist is still in playoff contention. We find that the passer's race coefficient is insignificant and similar in magnitude in both specifications, suggesting that the importance of the game itself does not appear to affect racial bias. Since this is a measure of the game's overall importance, our finding does not conflict with the in-game bias we observe as score differentials increase. These regressions are found in the online Appendix.

<sup>11</sup> Any adverse performance effects may still be reflected in their compensation, however.

<sup>12</sup> Analogously, Becker (1971) points out that employment discrimination is more likely to occur in firms with market power than in competitive firms. In competitive settings, non-discriminating firms will expand and drive out discriminating ones. In the case of a monopoly, however, a discriminating firm can exist indefinitely, though with lower profits.

averaged fewer than 10 minutes per game. We believe that these two groups represent inframarginal and marginal players respectively. In results not shown, we find that both of these groups exhibit insignificant amounts of own-race bias. Since in both groups, estimates are insignificantly different from zero, the two groups are insignificantly different from each other.<sup>13</sup>

### **Interpretation and Conclusion**

Assist patterns in the NBA exhibit very little evidence of same-race bias. More specifically, given a particular set of players on the court, a white passer is no more likely to pass to a white teammate than a black passer. While this result is interesting, it is important to note why the high degree of interracial cooperation may be specific to the NBA.

Our model can be thought of as the final node in a more complex game in which players are matched to teams, and coaches decide which players are on the floor and which plays to run. To the extent that some players exhibit strong same race preferences, they may be matched to teams with a high frequency of same-race teammates. With sufficiently high levels of racial segregation, the economic penalty of same race preferences could be quite small. Furthermore, conditional upon the team roster, coaches may choose personnel combinations in which efficient passing choices occur. Such optimizing behavior does not affect the consistency of our results. It does imply, however, that our results only reflect the level of same-race bias in situations that occur in equilibrium. The average level of same race bias may be higher.

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<sup>13</sup> The results are found in the Appendix online.

In many workplace environments, the effect of same-race bias may not be immediately apparent or have little impact on the actors involved. The NBA differs from such instances in that player behavior is closely observed by coaches, owners, and many thousands of fans, and the result of poor interracial cooperation may have an immediate effect on the outcome of the game. Also, to the extent that players derive utility from winning, they have an immediate incentive to engage in interracial cooperation.

Ultimately, our findings do not imply that efficient interracial cooperation occurs throughout the economy. They do imply, however, that interracial cooperation can occur in equilibrium when incentives are well aligned for efficient cooperation. Firms may want to consider how they can alter incentives to promote efficient cooperation among a diverse workforce.



**Table 1 - Summary Statistics for Full and Restricted Samples**

	All Observations		Black Passer		White Passer	
	Full Sample	Restricted Sample	Full Sample	Restricted Sample	Full Sample	Restricted Sample
	(1)	(2)	(3)	(4)	(5)	(6)
Black Shooter	0.7849 [0.4109]	0.6290 [0.4831]	0.7917 [0.4068]	0.6367 [0.4801]	0.7623 [0.4257]	0.5985 [0.4902]
White Shooter	0.2151 [0.4109]	0.3710 [0.4831]	0.2083 [0.4068]	0.3633 [0.4801]	0.2377 [0.4257]	0.4015 [0.4902]
Center	0.0651 [0.2468]	0.0593 [0.2406]	0.0477 [0.2132]	0.0438 [0.2046]	0.1339 [0.3406]	0.1226 [0.3280]
Forward	0.2797 [0.4488]	0.2584 [0.4378]	0.2729 [0.4455]	0.2529 [0.4347]	0.3063 [0.4609]	0.2808 [0.4494]
Guard	0.6552 [0.4753]	0.6823 [0.4656]	0.6793 [0.4667]	0.7033 [0.4568]	0.5598 [0.4964]	0.5965 [0.4906]
First year with team	0.3084 [0.4619]	0.2993 [0.4580]	0.3177 [0.4656]	0.3145 [0.4643]	0.2721 [0.4450]	0.2450 [0.4301]
N	302,714	149,059	241,613	118,321	61,101	31,741

*Notes:* Standard deviations in brackets. Information about position and first year with team refers to the passer. In the restricted sample, observations are dropped where there is no racial variation on the court.

**Table 2. Factors Associated with the Probability that the Shooter Is White (Using the Restricted Sample)**

	(1)	(2)	(3)	(4)	(5)
Passer is white	0.0353** (0.0122)	0.0026 (0.0061)	-0.0044 (0.0056)	0.0065 (0.0057)	0.0004 (0.0040)
Black players on the court besides the passer		-0.2178** (0.0036)	-0.2094** (0.0030)		
Prediction that shooter is white ( $\theta$ )					0.9029** (0.0076)
Fixed Effects	None	None	Team-year	Player-group	None
Observations	149,059	149,059	149,059	149,059	149,059
R <sup>2</sup>	0.0013	0.0729	0.0815	0.1524	0.0890

Notes: Standard errors (in brackets) are clustered at the team-year level. \* and \*\* indicate statistical significance at the 5\* and 1% levels respectively. The player-group fixed effects include a control for each unique combination of the four players on the court besides the shooter. Our measure for  $\theta$  comes from equation (3) on page 8. The sample is restricted to observations where there was at least one player of each race to receive the pass and to player-groups for which we observe at least one passer of each race in the data. Each regression includes controls for the passer's position and whether the passer is in his first year with the team.

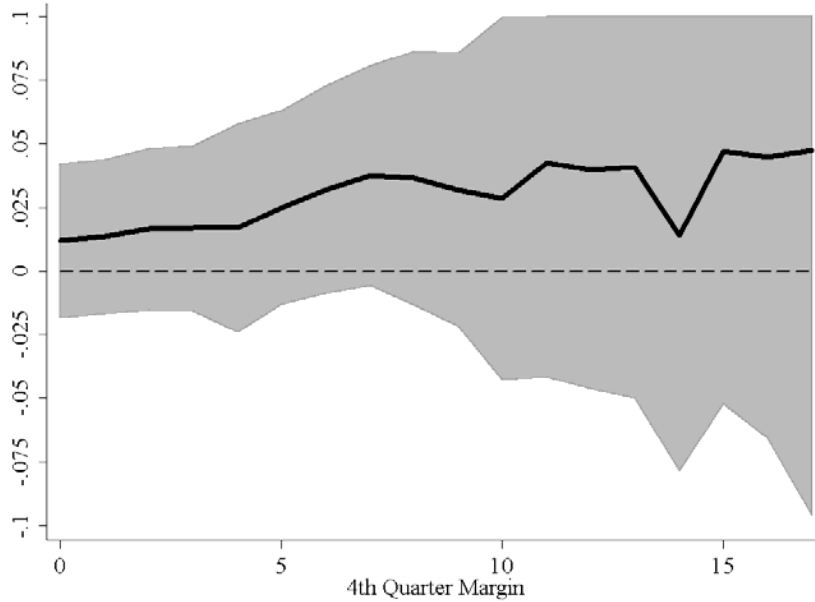
**Table 3. Factors Associated with the Probability that the Shooter Is White (Using the Full Sample)**

	(1)	(2)	(3)	(4)	(5)
Passer is white	0.0288 (0.0170)	0.0016 (0.0038)	-0.0023 (0.0036)	0.0035 (0.0033)	0.0002 (0.0022)
Black players on the court besides the passer		-0.2524** (0.0017)	-0.2456** (0.0022)		
Prediction that shooter is white ( $\theta$ )					1.0105** (0.0023)
Fixed Effects	None	None	Team-year	Player-group	None
Observations	302,714	302,714	302,714	302,714	302,714
R <sup>2</sup>	0.0008	0.2670	0.2721	0.4230	0.2847

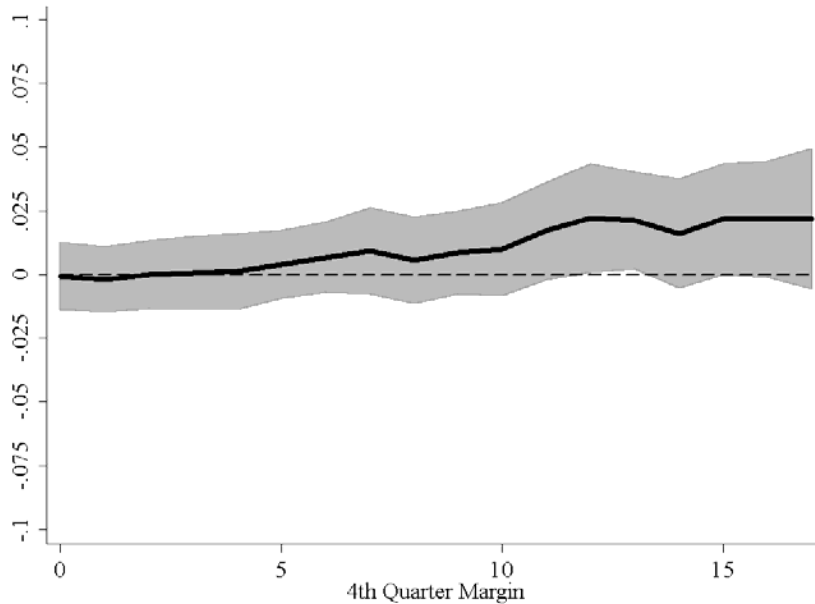
Notes: Standard errors (in brackets) are clustered at the team-year level. \* and \*\* indicate statistical significance at the 5\* and 1% levels respectively. The player-group fixed effects include a control for each unique combination of the four players on the court besides the shooter. Our measure for  $\theta$  comes from equation (3) on page 8. Each regression includes controls for the passer's position and whether the passer is in his first year with the team.

Figure 1. Change in estimated bias based on absolute value of score difference at the start of the 4<sup>th</sup> quarter.

A. Player-group fixed-effect estimate



B. Theta-proxy estimate



Notes: The 4<sup>th</sup> quarter margin refers to the absolute value of the difference in the two scores at the start of the fourth quarter. The dark line provides the estimated coefficient of the same-race bias for games in which the blow-out margin was at least as large as the number indicated on the x-axis. The shaded region provides the 95% confidence interval.

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