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Are Findings of Salary Discrimination Against Foreign-Born Players in the NBA Robust?*

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*We would like to acknowledge helpful comments and suggestions made by participants in a sports session at the 2015 Southern Economics Association conference in New Orleans, Louisiana.

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Abstract:

The influx of international players into the NBA has led researchers to investigate whether pay discrimination exists for these new entrants. Using a two-stage double fixed-effect model, Yang and Lin (2015) found evidence of salary discrimination against international players. Using a similar technique with a much longer unbalanced panel dataset (1989-2013), we are unable to verify their results.

Keywords: Wage Discrimination, NBA

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Proclamations of racial discrimination always elicit notoriety. Findings of no discrimination do not procure the same response. Therefore, it is important that any positive findings of racially unequal treatment be particularly robust. --Groothuis and Hill (2013)

The growing wave of international entrants into the National Basketball Association has led to recent research into the existence of pay discrimination against foreign players. Using yearly salary regressions Eschker, Perez, and Siegler (2004) suggested there was a premium paid to international players for the 1996-97 and 1997-98 seasons and posited a “winner’s curse” in the market due to an inability of NBA scouts and general managers to properly evaluate the worth of foreign players who did not play college basketball in the U.S. More recently using an unbalanced panel dataset (1999-2008) and a two-stage double fixed-effect model Yang and Lin (2015) find evidence of salary discrimination against international players.

The purpose of this paper is to check the robustness of the Yang and Lin (2015) results. Using a much larger unbalanced panel dataset (1989-2013) and the same econometric technique, we are unable to verify the existence of pay discrimination against foreign players. In addition, even when we truncated our data and use the same panel study as Yang and Lin (2015), we find that their results are not robust to various specifications. In the next section we explain the econometric technique and then discuss our dataset and the differences between the variables included in our regressions compared to Yang and Lin (2015) regressions. Then we report our empirical results and compare them to the Yang and Lin (2015) findings. Lastly, we conclude with a discussion concerning the robustness of pay discrimination research.

Econometric Technique

We use the two-stage fixed effects model developed by Bartel and Sicherman (1999) and utilized by Yang and Lin (2015). In the first stage standard salary estimation equations utilize the natural logarithm of salary as the dependent variable and a vector of productivity measures as the independent variables in a fixed effect panel. In a panel setting this vector can be classified into two subsets, one which changes yearly with performance (X_{it}) and one which does not change yearly but may affect productivity (Y_i). Thus our equation can be written as:

$$\ln \mathrm{S}_{\mathrm{it}} = X_{\mathrm{it}}\alpha + Y_i\beta + u_i + e_{it}\,. \qquad (1)$$

In a panel setting it is assumed that the vector Y_i consists of unchanging characteristics of each individual and therefore only variables found in vector X_{it} are included in the initial fixed-effect regression. Since the panel is unbalanced the standard errors are clustered. The residuals resulting from this fixed-effect regression should reflect the unobserved fixed component of the error term attributable to time-invariant differences between individuals. Bartel and Sicherman (1999) identify this as an individual wage "premia." The premia is the fixed component of the salary that is not explained by time varying performance characteristics. This premia is individually fixed and can be either positive or negative.

Following Bartel and Sicherman (1999) and Yang and Lin (2015) these residuals or wage "premia"s are then regressed on the unchanging individual characteristics of individuals, vector Y_i. Given the unbalanced nature of the panel this regression is estimated using weighted least squares. The weights used by the Stata program are inversely proportional to the variance of an observation. This equation takes the format:

$$\hat{u}_{\mathrm{i}} = Y_i\beta + \omega_{\mathrm{i}}. \quad (2)$$

The regression on the individual wage "pemia"s can then identify the influence of time invariant characteristics such as height, race, draft number or being foreign born. One key component of the technique is that the wage "premia" is the residual from stage one regressions; therefore the stage one specification is crucial for the stage two results.

Dataset and Variables

The dataset for our analysis includes all players in the NBA from the 1989-90 season thru the 2012-13 season. Variables used to measure on-court performance include experience, experience squared, points scored per game, rebounds per game, assists per game, steals per game, turnovers per game, blocks per game, games played, and minutes per game. Except for turnovers all of these measures are also included in the Yang and Lin (2015) models. Unlike Yang and Lin (2015) however, dummy variables for player positions are not included in the fixed-effects regression. Player positions rarely, if ever, change from season to season. Inclusion of these dummies is inappropriate in the fixed-effect model because the time invariant variable of height captures the same effect. We therefore estimate equation 1 with the variables listed above as an unbalanced panel using a fixed-effect model and clustered standard errors.

For the estimation of equation 2 on each individual's wage 'premia', we use the independent variables of time-invariant personal player characteristics that includes the player's height, a dummy variable for white players, years of college, a dummy variable if the player was drafted in the first round, the player's overall draft number, a dummy variable for players born outside the U. S., a dummy variable for players born outside the U.S. who did not play college basketball in the U. S., and lastly a dummy variable for players born outside the U. S. who did play college basketball in the U.S. Our list includes all the variables used by Yang and Lin (2015) except for a measure of the domestic market of the player's team. However, unlike Yang and Lin (2015) we use three different classifications for foreign-born players. Eschker, Perez, and Siegler (2004) suggested the wage premium for foreign-born players in the NBA for 1996-97 and 1997-98 seasons was only true for those players who did not play in college in the U.S. Therefore the current research will designate each classification separately as well as collectively.

Tables 1 and 1a show the means and standard deviations of the variables used in equations 1 and 2. These statistics are shown by the number of observations and by the number of players. Some interesting insight can be gleaned from this data. First, on average, foreign-born players are taller by two inches or more. Second, the majority of foreign-born players who did not play basketball in college in the U.S. are white. This is not true for foreign-born players who did play college basketball in the U.S.; only around 40% of these players are white. Third, foreign-born players in all categories earned higher real salaries on average compared to their counterparts born in the U.S. This is true despite that fact that foreign-born players are generally taken lower in the draft on average; the only exception to this statement is for foreign-born players who played college basketball in the U.S. and were first round draft picks. In terms of performance statistics, foreign-born players show better rebounding and block statistics compared to native-born players; obviously this stems from their greater height.

Empirical Results

We report the results for the regression models for equation 1 in Table 2. Six different versions of the model are shown. It is expected that the coefficient of experience is positive and the coefficient of experienced squared is negative; this reflects the standard experience/earnings profile shape. Indeed, all six versions of the models show correctly signed and significant coefficients at the 1% level. In all three specifications including age and age squared we find that the coefficient on age is negative and significant while age squared is insignificant. Given the high degree of multicollinarity between age and experience we suggest that the negative coefficient on age captures the likelihood of the players who enter the NBA at an older age are less skilled then early entrants as found by Groothuis et al.(2009). When it comes to performance, it is expected that rebounds per game, assists per game, blocks per game, steals per game, points per game, games played, and minutes per game should have positive coefficients. Only models 2 and 5 show correctly signed and significant coefficients for the variables included. In other models some of the performance variables are incorrectly signed and significant. This illustrates the high degree of collinearity between these measures. For example, guards handle the ball a great deal of the time in the NBA. Therefore they are likely to have the most assists on the team. However, this also makes them likely to have the most turnovers. Tall players who rebound well are also more likely to have more blocks shots. Lastly, including minutes played per game causes many of the performance variables to lose explanatory power since only the players who contribute the most, play the most. Our findings on the performance variables are similar to those of Yang and Lin (2015).

To estimate equation 2 the individual wage "premia"s or residuals from each of the equations in Table 2 are regressed on time-invariant personal characteristics using the same weighted least squares approach used by Yang and Lin (2015). Unlike Yang and Lin (2015) various different models are explored for the second stage WLS regressions including three different categories for foreign-born players. Another major difference between the results in Table 3 and those of Yang and Lin (2015) and Eschker, Perez, and Siegler (2004) is the inclusion of a player's overall draft number in some versions of the model as opposed to just a dummy variable for first round draft picks. The NBA has institutionalized salaries for rookie scale contracts for first-round picks and the differences are dramatic. For example, the first pick in the 2012 draft received a salary for the 2012-13 season of $5,144,280 as listed in the collective bargaining agreement between the league and union whereas the last pick in the first round received a salary of $1,020,960. This differential persists for the first three year of the rookie scale contracts.

In the interest of space only the WLS regressions using the residuals from models 1 and 4 in Table2 are shown in Table 3.[1] A few things are clear from an overview of these models. First, the dummy variable for first round draft picks is always positive and significant at the 1% level in models in which it appears. Similarly, the overall draft number is negative and significant at the 1% level in all models in which it appears. The coefficient of height is positive and significant at the 1% level in all models. The coefficient of years of college is negative and significant in six of the eight specifications in which it appears. . The dummy variable for white players is only significant in model 9; it is negative and significant at the 5% level.

Interestingly, we find that in all specifications based on residuals from Model 1 in Table 2 (age and age squared included), the coefficient on foreign born is positive and significant suggesting foreign born players are paid a premium which is the opposite of the Yang and Lin (2015) results. We find that the foreign born players who do not play college basketball in the U.S. receive the highest premium. However, for models 6 thru 10 based on Model 4 in Table 2 (no age and age squared),model 9 is the only model in which any dummy variable for foreign-born players is significant; the coefficient for players born outside the U.S. who did not play college basketball in the U.S. is positive and significant at the 1% level. The only difference between models 9 and 10 is the addition of years of college as an explanatory variable. When it is added to the model the coefficient for players born outside the U.S. who did not play college in the U.S. becomes insignificant.

The results in Table 3 demonstrate robust relationships between the residuals of the fixed-effect salary regressions and height and draft status. The results for the other variables in the model are mixed. It is obvious from the various models that there is collinearity between the white dummy variable and the dummy for players born outside the U.S. without U.S. college experience. It is also clear that years of college and the dummy for players born outside the U.S. without U.S. college experience are collinear since all of these foreign-born players have no college experience. Overall our results do not provide support for the Yang and Lin (2015) findings of pay discrimination against foreign-born players no matter what the classification[2].

Conclusions and Implications

Economics and econometric analysis have provided useful insights into the presence of pay discrimination in this country and helped to shape policy measures designed to alleviate such discrimination. Given that findings of racial discrimination elicit notoriety and that findings of no discrimination generally do not procure the same response it is important that any positive findings of racially unequal treatment be particularly robust. . Unfortunately the finding of pay discrimination against foreign-born players in the NBA found by Yang and Lin (2015) does not meet this standard.

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