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Do Major League Baseball Hitters Engage in Opportunistic Behavior?

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International Advances in Economic Research
DO MAJOR LEAGUE BASEBALL HITTERS ENGAGE IN OPPORTUNISTIC
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DO MAJOR LEAGUE BASEBALL HITTERS ENGAGE IN OPPORTUNISTIC BEHAVIOR?

Abstract This study focuses on 256 Major League Baseball free agent hitters playing under the 2006-2011 Collective Bargaining Agreement to determine whether players engage in opportunistic behavior in their contract year, i.e., the last year of their current guaranteed contracts. Past studies of professional baseball yield conflicting results depending on the econometric technique applied and choice of performance measure. When testing whether players' offensive performances increase during their contract year the omitted variable bias associated with OLS and pooled OLS estimation lead to contrary results compared to fixed effects modeling. Fixed effects results suggest players increase their offensive performance subject to controlling for the intention to retire.

JEL Classification Codes: J31, J44

Keywords: sports economics, free agency, opportunistic behavior, contract year, guaranteed contracts

Introduction

In sports, especially baseball, there is a lot of talk about contract year performance. Beginning in spring training and continuing throughout the season, sports journalists and fans converse about how players in the last year of their contract will perform that season. Many followers speculate that contract year players will have break-out seasons in order to secure a better contract in upcoming contract negotiations. Others contend the anxiety over the need to procure a contract for the following season distracts players, causing them to underperform in their contract year. Baseball Reference.com (2012) shows Jason Werth playing for the Phillies for \$7.5 million in 2010. Being the last year of his contract, Werth sported a .921 on base plus slugging percentage (OPS) compared to .879 in the previous year. Following his stellar 2010 contract year season, Werth signed a seven-year contract with the Washington Nationals worth \$122 million or more than \$17 million per season. Werth's performance differs from the Phillies' Pat Burrell who had a .902 OPS in 2007, which dropped to .875 during his contract year in 2008. Rather than speculate the contract year phenomenon with anecdotes, sufficient sample sizes and appropriate econometric analysis provide robust evidence.

If players increase their effort and performance during their contract year, it implies potential opportunistic behavior by players, a topic studied by labor economists examining performance and incentives in contracts. Utility-maximizing players face profit-maximizing team owners when negotiating contracts. Both sides understand incentives affect performance, and performance impacts pay. Understanding opportunistic behavior and whether players engage in it aids team owners in crafting contracts. The labor market for sports teams offers empirically

robust economic data to investigate opportunistic behavior since publicly available player performance indicators, contract clauses, and salary compensation are available. Not surprisingly, using different theories, variables, data, or time periods lead to a variety of results.

This paper focuses on position players (non-pitchers) in baseball in order to isolate individual performance as opposed to performance dependent upon teammates. Pitching performance is more greatly affected by team defensive capabilities, thus excluded in the analysis. For hitters, the type of pitch a hitter faces may depend in part on the hitters adjacent to him in the line-up, but this is dismissed as a reason for changes in batting performance from one year to the next because players tend to bat in similar slots in the line-up across seasons. Among the various offensive performance measures used in other contract year studies, the statistics on-base percentage plus slugging percentage (OPS) and adjusted by ballpark on-base-percentage (OPS100) are chosen for this study as they measure a hitter's output independently from his teammates' contributions.

The salary structure in Major League Baseball (MLB) emanates from the collective bargaining agreement (CBA) between the 30 MLB teams and the Major League Baseball Players Association (MLBPA). The most recently completed CBA dictated employment conditions from December 26, 2006 until the end of 2011 season. That CBA set the minimum MLB salary at \$400,000 in 2006, rising to \$440,000 by 2011, and set the minimum minor league salary for players with MLB experience at \$65,000 (Bloom 2006). The average salary, however, in the MLB in 2011 was \$3,095,183, demonstrating the tremendous boost in income above the minimum pay by playing in the majors (MLBPA Info 2012). As is the case with most jobs, pay is determined by performance in MLB, so that players are motivated to perform better to increase their pay. Unlike most jobs, however, current contracts in MLB are guaranteed and teams are obligated to pay the agreed amount no matter how the player performs during the contract's length. Thus, there is a large financial incentive to graduate from the minor leagues, land a MLB guaranteed contract, and not be demoted back to the minors. Moreover, the income associated with performance-based incentive clauses in contracts, such as earning an All-Star spot or league MVP, are small relative to yearly earnings. For example, in 2012, Ryan Howard would have received \$25,000 if he made the All-Star team and \$100,000 as the National League MVP compared to his \$20 million salary (Baseball Reference.com 2012).

Making it to the majors, however, can be just the beginning of high salaries. Once a player signs his first contract with a team, he accepts the team's salary offer for his first two years of service. He can potentially increase his initial salary through arbitration. A player with three or more years of MLB experience, but not more than six years, is eligible for salary arbitration. Superlative players following their second season may be arbitration-eligible if they meet certain service thresholds. Salary arbitration revolves around high versus low salary offers. In January of a given year, a player eligible for arbitration and his team both submit a proposed salary to a three-person panel of professional arbitrators, and a final-offer hearing takes place

between the 1st and the 20th of February. If the player and team do not agree on a salary prior to the arbitration panel's decision, the panel's choice becomes the new one-year guaranteed salary. In deciding whether to award the player the higher or lower salary offer, the panel considers the following criteria: the player's contribution to the team in terms of performance and leadership, the team's record and attendance, awards won by the player, all-star games in which the player appears, postseason performance, and the salaries of comparable players with the same service-time. The two sides cannot talk about the team's finances or previous negotiations (Ray 2008). For example, following a superb 47 home run-season in 2007, The Phillies' Ryan Howard asked for \$10 million in arbitration and the Phillies countered with a \$7 million offer. Having just completed his second season with the Phillies, Howard won arbitration and increased his 2008 salary to \$10 million, a substantial increase from his 2007 salary of \$900,000 (Stark 2008).

Besides arbitration, there are two other avenues for a player to seek more money and more years on his contract. First, a player can renegotiate with his team to change his current contract before expiration of the contract, i.e., seek a contract extension. In February 2009, as a World Series winner and runner-up in the National League MVP Award, Ryan Howard with three years of arbitration eligibility remaining with the Phillies signed a 3-year, \$54 million contract. In 2010, Howard extended his contract again to five years for \$125 million (Baseball-Reference 2012). Alternatively, players become eligible for free agency with the ability to shop their talents to any team.

Players with six seasons of service may enter free agency. Additionally, players who have at least three years of MLB experience and are sent down to the minor leagues also have the option to decline this demotion and enter free agency. Players declaring free agency must do so within the 15 day period beginning on October 15th or the day after the last game of the World Series, whichever is later.

MLB's salary arbitration, contract extensions and free agency all provide avenues for better contract conditions for players. Players' past and current performances serve as predictors for future performance, which team owners scrutinize when negotiating new contracts. With so much money at stake, are players more likely to boost their effort and performance to secure their financial future? While opportunistic behavior may occur during salary arbitration-eligible years and players seeking renegotiated contracts, this paper focuses only on MLB free agents with six or more years of service for four reasons: free agency is associated with the greatest financial gains for players as teams bid for players' services due to the declining monopsonistic power of owners after six years; at least six years of service enables more observations per player to capture more robust results; free agents with fewer than six years are those who have been demoted to the minors; and there will be a sufficient number of players who intend to retire at the end of their contract year, an intention that is expected to impact contract year performance.

Previous Research Findings

The majority of previous research testing for contract year performance examines the MLB (Harder 1991; Kahn 1993; Dinerstein 2007; Maxcy et al 2002; Krautmann and Oppenheimer 2009; Martin et al 2011) and the results vary. Stiroh (2007) studies the National Basketball Association (NBA), which also guarantees player contracts. We learn from the studies that the unit of measurement for performance and econometric technique employed are often the sources of contradictory results.

If performance varies from year to year, one explanation is that effort has changed from one year to the next due to performance-enhancing incentives in one year compared to another. Alternatively, players may exhibit random fluctuations around an unknown average performance level. Albert and Bennett (2001) find a .088-point range in on-base percentage (OBP) when simulating 100 seasons for a player with an average OBP of .380. While random fluctuation is feasible, it is unlikely that above average performances occur randomly during the contract years for a large group of players.

Maxcy et al (2002) develop several performance measures using the deviation between the current year and a three-year moving average of particular hitting or pitching statistics. For hitters, they use the difference in current year slugging percentage (SLG) and the 3-year moving average of the hitter's SLG as a dependent variable to measure a hitter's deviation in offensive performance, while controlling for skill levels across players. Controlling for age and field position, they find no statistically significant increase in hitters' performances during the last year of their contracts. However, using deviations in playing time and durability, they find hitters spend 4.9 fewer days on the disabled list and have 24.28 more at bats during the contract year. The choice of dependent variable matters in testing for opportunistic behavior.

Other hitting performance measures, Wins Above Replacement Player (WARP) and Runs Created per 27 outs (R27), are used by Perry (2006) and Birnbaum (2002), respectively. WARP incorporates offensive and defensive capabilities, and the number of games played. Perry (2006) discovers evidence of boosted contract year performance for 212 top free agents from 1976-2000. Fisher (2009), however, points out that managers dictate playing time so that players have less control over their WARP than their OPS, and contends OPS is the preferred offensive statistic. Birnbaum (2002), using R27, does not find evidence of higher contract year performance. Fisher (2009) contends Birnbaum's results may be due to failing to eliminate players with too few playing performances, leading to unstable R27 values.

The choice of the econometric technique also affects results. Martin et al (2011) study 160 randomly chosen free agent hitters from 1996-2008 for five offensive performance measures and fail to find the contract year phenomenon using one-way analysis of variance (ANOVA) techniques. They limit the players to those with multiyear contracts and at least 250 at-bats per year, which may reduce the number of retiring players in the sample. Maxcy et al (2002) use ordinary least squares (OLS) estimation on their 213 hitters with more than three years

experience and multiple observations per hitter, which leads to a 1,160 sample size. OLS may not provide the best econometric method available for panel data since it fails to address omitted variable bias. Kahn (1993) addresses this concern by using a fixed effect regression model with his 1987-1990 longitudinal data of MLB hitters and pitchers when estimating salary arbitration and free agency effects on salaries and contract lengths. Kahn finds free agency positively affects salary and contract length, two desired outcomes for players.

As CBAs change over time, newer data sets are needed to re-examine opportunistic behavior. Dinerstein (2007) employs 1,330 hitter-years from 2001-04. He not only examines contract year performance by hitters, but whether teams create incentives that lead to contract year behavior. Using seemingly unrelated regression (SUR), he observes the contract year increases SLG by .0095 for an equivalent of \$138,813 on a free agent's contract. Interestingly, from the teams' perspectives, when making a salary offer to a free agent, consistency in a player's SLG mattered more than the most recent SLG. This finding suggests teams are not creating the contract year incentive and that players acting rationally will aim for hitting consistency, *ceteris paribus*. As players digest this undervalued contract year effect, they may aim for steady hitting performances.

Greater contract year performance may also be due to expectancy theory, which entails two expectations: a player expects greater performance will lead to a desired outcome (higher salary and/or longer contract length) and increased effort is expected to enhance performance. This implies the free-agent hitter understands which offensive measure is associated with a better contract. Harder (1991) estimates OLS models on data from 106 hitters (17 free agents) from 1977-1980 and finds home runs were more generously rewarded than batting average, yet no evidence of increased home run production in the contract year. He attributes the statistically significant decline in batting average during the contract year to equity theory, that hitters felt underappreciated with high batting averages and therefore reduced their effort therein. Ahlstrom et al (1999) use data on 172 free-agent hitters from 1976-1992 who changed teams and for five offensive statistics find no statistically significant improvements in performance in the contract year.

Knowing the link between performance and pay is necessary before any opportunistic behavior can occur. Stiroh (2007) uses a composite performance rating that captures points, rebounds, assists and blocks for 263 NBA players to test for contract year performance. He first shows that a one point increase in the historical and contract year composite performance ratings increase salaries by 13% and 8%, respectively. Second, given the 8% contract year incentive, he uses 2,077 player-years for 349 players with multi-year contracts from 1988-2000 with fixed effects estimation and finds evidence of improved performance during the contract year.

Economic Theory

The economic theory behind the contract year phenomenon is well stated by Stiroh (2007). A moral hazard exists given asymmetric information between the agent (player) and the principal (team owner) due to the uncertainty of a player's future performance. Performance, demanded by owners, depends on a player's effort and ability. Ability differs across players and is not easily measured or distinguishable from effort. A player's ability is assumed to be relatively constant over time, but effort can change from season to season leading to different performance levels. As a professional athlete, baseball players are constantly judged on their performance by their statistics, so it is unlikely effort changes much during a game. However, off-season effort and effort between games during the season can vary. These efforts are largely unobservable and are not subject to scrutiny of the public unless the media uncovers aberrant behavior. Owners can see differences in performances across players, but it is not clear how much is due to innate ability or effort intensity.

Assuming ability is predetermined, players choose to exert an amount of effort to maximize their utility by equating the marginal benefit of higher income and team success derived from effort to the marginal cost of less leisure commensurate with effort. Meanwhile, team owners craft contracts to maximize profits by paying players according to their performance, which yield wins and revenues for the team.

With guaranteed contracts, the disconnect between contemporaneous performance and pay leads to a moral hazard in that poor performance after signing a contract does not lead to a pay cut. This makes it imperative that owners weigh players' past performances and players' other traits wisely when predicting future productivity, and that they try to segregate effort from ability. Players, on the other hand, seeking a series of contracts prior to retiring, will choose effort to maximize their present discounted utility of their current and future contracts. Players alter their effort according to how they perceive owners will reward them. If players believe owners overweight a player's most recent performance, this incents the player to bolster his contract year performance to ensure a desirable future contract. Hence, players increase their effort and performance during their contract year. If, however, general managers value consistency by weighing all past performances equally when crafting a new contract and a player perceives this, then boosting performance in the contract year is unlikely. Thus the player's effort and performance depend on his perception of the owner's algorithm for using past performance to predict future performance.

When a player intends to retire at the end of the contract cycle, the incentive to perform to gain another contract is gone. Intended retirement is expected to reduce effort and performance during all years of the final contract, including the last year. Effort and performance during the "contract year" for a retiring player will not be expected to be as high as a player desiring another contract.

Based on the above, two testable hypotheses follow. First, if the most recent performance yields a highly desired contract, meaning the contract year incentive exists, one expects to witness such opportunistic behavior. Second, the contract year performance for players intending to retire will be less than those who plan to keep playing.

Population Regression Function

On base percentage measures a player's ability to get on base not due to a fielding error relative to his official at-bats. A player reaches base by getting a hit, drawing a walk or being hit by a pitch. Getting on base is necessary to ultimately score an earned run. Slugging percentage measures power hitting because each additional base earned is weighted more, meaning home runs have a weight of four compared to singles with weights of one. The more bases achieved during an at-bat, the greater the likelihood of more earned runs scored. Summing OBP and SLG to create OPS yields a diverse offensive measure because it accounts for both power and reaching base frequently. A player with a high OPS reaches base often and hits extra-base hits, which contribute to his team's runs scored. Barry Bonds is the single-season record holder for OPS at 1.4217 in 2004. His SLG was .812 and his OBP was .609 (Baseball-Reference 2012). During that season he was typically walked or hit a home run during a plate appearance, explaining why he had such a high OPS.

The dependent variables for this study are OPS and OPS100 for hitters. OPS100, the adjusted OPS, accounts for league and home baseball park in which the player plays and is standardized to a league average of 100. Keri (2006) contends OPS ably measures a player's offensive abilities more than hitting components such as RBIs, at bats (AB), batting average (BA) and home runs (HR). Also, OPS is not dependent upon playing time, which is determined by the manager. Albert and Bennett (2001) find OPS is a better predictor of scoring runs, the chief goal of a team's offense, than its two components OBP and SLG separately.

The population regression model for the OPS (or OPS100) for player i in season t is

$$\begin{array}{ccccccc}
 & (+) & & (-) & & (?) & & (?) \\
 (1) \text{ OPS}_{i,t} = & \beta_0 + & \beta_1 * \text{AGE}_{i,t} + & \beta_2 * \text{AGE2}_{i,t} + & \beta_3 * \text{GAMES}_{i,t} + & \beta_4 * \text{POSITION}_{i,t} \\
 & (+) & & (-) & & (+) & & \\
 & + \beta_5 * \text{PLAYOFF}_{i,t} + & \beta_6 * \text{RETIRE}_{i,t} + & \beta_7 * \text{CONTRACTYR}_{i,t} + & a_i + v_t + \mu_{i,t},
 \end{array}$$

where AGE is the player's age; AGE2 is the square of a player's age; GAMES is the number of games in which the player played; PLAYOFF is a binary variable equal to 1 when the player's team makes the playoffs and 0 otherwise; POSITION uses dummy variables for the various fielding positions; RETIRE is a dummy variable equal to 1 for a retiring player and 0 if not; and

CONTRACTYR is a dummy variable denoting whether season t was a contract year ($=1$) or not ($=0$). The signs above each coefficient denote the partial derivatives' expected impact on OPS.

The stochastic error comprises three terms that impact a player's OPS: a_i is the unobserved player effect representing all time-invariant factors that cannot be measured or observed, such as innate ability and family background; v_t represents immeasurable or unobserved time-variant player characteristics, such as risk tolerance and personal/familial events; and μ_{it} is the stochastic error term.

The model uses a quadratic model to represent the age of the player during free agency on hitting performance. AGE2 is expected to have a negative coefficient, implying OPS increases at a decreasing rate as a player ages until reaching a threshold, then falls due to the depreciation of the player's body from wear and tear. The quadratic impact of age on productivity shows a threshold near age 39 in Maxcy et al (2002). Krautmann and Oppenheimer (2002), followed by Link and Yosifov (2012), use years of experience to capture how years of playing the game in the majors affects performance. These researchers, however, exclude the quadratic term for experience due to high degrees of multicollinearity. Krautmann and Solow (2009) estimate the experience threshold is about 11 years for free agents.

Two variables are used as control variables to mitigate potential bias. The first, GAMES, accounts for playing time and its predicted sign is unclear. An oft-injured good player plays in fewer games but can still have a high OPS. However, playing more games may help a player gain confidence behind the plate and raise his OPS. The second, POSITION, captures the plurality field position played in a season. Grouping positions is also possible. For example, Krautmann and Solow (2009) use catchers and shortstops as a control group relative to all other fielding positions. They would expect $\beta_4 < 0$ because catchers and shortstops are relied on for their defensive prowess more than their hitting abilities, thus have lower OPS's, *ceteris paribus*.

The expected sign on PLAYOFF is positive. If a player's team makes the playoffs, one expects an increase in his production because playing for a championship gives a player an incentive to perform better. Also, playoff teams pay their players a share of the playoff revenue, providing a financial incentive to perform better. It is also possible that teams in the playoff hunt trade for high performing hitters at the trade deadline. For players who change teams during the season via a trade, the playoff status of the final team upon which the player played is used.

An appropriate functional form necessitates the inclusion of a retirement variable. O'Neill and Hummel (2011) model retirement as an ex-post condition wherein a player is considered retired if not playing in the major league the year after the contract year. The problem, of course, is it is not clear whether the player did not return the following year because no team wanted his services and yet he wanted to continue to play, or he truly wished to retire. Their ex-post retirement variable indicates an 11.2-11.3% decrease in OPS, *ceteris paribus*. The

appropriate retirement variable should be ex-ante, since the intention to retire is expected to decrease the contract year performance due to the lack of incentive to garner another contract. Lacking an explicit intentionality measure, Krautmann and Solow (2009) estimate the probability of retiring as a function of the player's age, OPS100, and days on the disabled list during the contract year. Their dependent variable is 1 when the player did not return to the majors the year following the contract year and 0 if he did return, i.e., did not retire. Although their likelihood function still uses the ex-post observation for retirement, their predicted retirement variable reduces the bias associated with a simple ex-post measure. The expected sign on β_6 is negative because players who expect to retire, or are more likely to retire, or exhibit ex-post retirement, are more likely to have a decline in their offensive performance because they do not have the incentive of signing another contract.

Per the marginal cost and benefits of effort discussed by Stiroh (2007), MLB hitters engage in opportunistic behavior and increase their performance during the contract year, thus $\beta_7 > 0$. This presumes team owners value the most recent performance as a solid indicator of future performance, which is why this opportunistic behavior occurs.

Data

Data were collected on all free agent hitters playing during the most recently completed 2006-2011 CBA with six or more years of MLB experience and with a minimum of two years of observation. By only choosing players under the same CBA, all players and team owners were subject to the same contract and free agency guidelines and operated with the same revenue-sharing rules. Doing so eliminates any potential impacts due to changes in the CBA. Hitters with one-year contracts or longer term-contracts are used. As Kahn (1993) and Stiroh (2007) show, players with longer term contracts are generally those with higher ability, so that eliminating those with one-year contracts would potentially bias the results by focusing on high ability players. Ultimately, 256 MLB free agent hitters met the data selection requirements.

ESPN.com's MLB Baseball Free Agent Tracker lists the positions played, age, current team and new team, if re-signed, for all free agents in each year. Players who do not receive another contract are listed as retired or free agent again. The Baseball Reference.com website provides the OPS and OPS100 offensive performance measures, the number of games played each season, and the year in which a player debuted in the major leagues.

Short of interviewing each player, it is difficult to find a player's intention to retire at the end of a season. Few players tip their hand prior to or early in the season to say they are playing their last season. Three different variables to try to capture retirement intention are possible. The first is the ex-post retirement variable used by O'Neill and Hummel (2011). Players who do not appear on a MLB roster the year following their contract year, which is apparent from the Baseball Reference.com website, are noted as NOPLAY=1. Players still on rosters show

NOPLAY=0. Alternatively, reading newspaper articles and blogs about ex-post retirees can lead to discerning whether the players intended to retire. When evidence of retirement intention exists, RETIRE = 1. An example is Mike Redmond who was quoted in the Star Tribune saying he knew the 2010 season would most likely be his last (Christensen, 2010). RETIRE = 1 suffers from measurement error because the author's discretion is used when the evidence is murky. It is more apparent if retirement is not intended. If the ex-post retiree continued to file for free agency but was not chosen by any MLB team or offered a minor league contract, then there is no intention to retire, i.e., RETIRE=0. Obviously, players still on MLB rosters show RETIRE=0. Third, PROBRET, a likelihood of retirement variable, is created akin to Krautmann and Solow (2009). Josh Hermismeyer has provided the author with the number of days on the disabled list (DL) for all players in 2006-2009 from his MLB Injury Report. BackseatFan.com and Fansgraph.com provide the days on the disabled list for 2010 and 2011 players, respectively. These DL days are used to create the predicted retirement variable for each free agent. The econometric modeling creating PROBRET follows later in this paper.

The panel dataset is unbalanced because the years played in the majors between 2006 and 2011 are not the same for all the free agents. Table 1 presents the format of the unbalanced data set for two players. The first player is outfielder (POSITION=9) Bobby Abreu, given an identification code of 2, who played all six years of the 2006-2011 CBA and in 2102 played with the Dodgers, thus NOPLAY = 0 and RETIRE = 0 for all six years. In his 2008 contract year and the prior year, he was a member of the Yankees, having been traded from the Phillies to the Yankees in 2006. The Yankees made the playoffs in 2006 and 2007 but not in 2008, as shown by Playoff = 1 and 0, respectively. Abreu had an OPS of .814, an OPS100 of 113, and played in 158 games in 2007, compared to an OPS of .843 and an OPS100 of 120 in 156 games in 2008. Abreu debuted in the majors in 1996, implying 11 years of experience by 2006 and never appeared on the disabled list over the six years.

The second player, outfielder Moises Alou, shows two contract years, one in 2006 with the Giants and the other in 2008 with the Mets. He had 16 years of MLB experience by 2006 at age 39. Alou did not play on a MLB team in 2009, thus NOPLAY = 1 for 2008. Two major injuries in 2008 limited his playing time to only 15 games. According to a reporter from the Washington Post (Smith 2009), Alou did not wish to play on a part-time basis, and given his injuries, e.g., 163 days on the disabled list in 2008, he claimed he would rather retire than play part-time. He announced he would retire following the World Baseball Classic in early 2009, which is why RETIRE=1. Although Alou did not announce his intention to retire prior to or during the 2008 season, his later remarks about injuries and retiring are retrofitted to RETIRE=1 at the author's discretion.

Table 1: Unbalanced Dataset Example

NAME	CODE	YR	YEAR	TEAM	AGE	CONTRACT YEAR	OPS	OPS100	GAMES	POSITION	RETIRE	NOPLAY	DEBUT	EXP	PLAYOFF	DL
Bobby Abreu	2	1	2006	Phil/Yan	32	0	0.896	126	156	9	0	0	1996	11	1	0
Bobby Abreu	2	2	2007	Yan	33	0	0.814	113	158	9	0	0	1996	12	1	0
Bobby Abreu	2	3	2008	Yan	34	1	0.843	120	156	9	0	0	1996	13	0	0
Bobby Abreu	2	4	2009	Ang	35	0	0.825	118	152	9	0	0	1996	14	1	0
Bobby Abreu	2	5	2010	Ang	36	0	0.787	118	154	9	0	0	1996	15	0	0
Bobby Abreu	2	6	2011	Ang	37	0	0.717	104	142	9	0	0	1996	16	0	0
Moises Alou	4	1	2006	Giant	39	1	0.923	132	98	9	0	0	1990	17	0	50
Moises Alou	4	2	2007	Mets	40	0	0.916	137	87	9	0	0	1990	18	0	76
Moises Alou	4	3	2008	Mets	41	1	0.777	107	15	9	1	1	1990	19	0	163

Sorting the descriptive statistics by contract year status, interesting results appear, as shown in Table 2. The differences in means for all variables, except playoffs and days on the disabled list, are statistically significantly different at $p < .001$. There are 546 player-year observations for contract years and 470 for non-contract years. The average OPS for the contract year is .70 compared to .752 for the non-contract year, which is contrary to expectations if one believes greater performance occurs during the contract year. O'Neill and Hummel (2011) find a similar contrary expectation where the average OPS for the contract year is .722 compared to .745 for the non-contract year. If one were to ignore multiple regression analysis and rely on differences in means, per Holden and Sommers (2005), it would suggest players underperform in their contract year. O'Neill and Hummel (2011) explain the contrary expectations are due to the influence of ex-post retiring players pulling down the average OPS for contract year. Table II shows this relatively large number of ex-post retirements for contract year observations at 23.1% compared to only 3.2% for non-contract year cases. Maxcy et al (2002) show more at-bats for the contract years. Using games played as another measure of playing time, Table 2 suggests a contrary finding to Maxcy et al (2002). Non-contract year cases show an average 115.5 games played compared to 96.2 games for contract years.

Table 2: Descriptive Statistics for Contract Year versus Non-Contract Year

	CONTRACT YEAR					NON-CONTRACT YEAR				
	N	MEAN	ST. DEV.	MINIMUM	MAXIMUM	N	MEAN	ST. DEV.	MINIMUM	MAXIMUM
OPS	546	0.7	0.12	0.284	1.067	470	0.752	0.12	0.222	1.114
OPS100	546	85.899	30.41	-21	182	470	97.2	29.87	-39	192
RETIRE	546	0.086	0.28	0	1	470	0.008	0.09	0	1
NOPLAY	546	0.231	0.42	0	1	470	0.032	0.18	0	1
AGE	546	33.59	3.28	26	48	470	32.25	3.02	24	47
YEAR SEXP	546	11.6	3.29	7	26	470	10.6	3	6	25
PLAYOFF	546	0.333	0.47	0	1	470	0.309	0.46	0	1
DL	546	19.39	33.74	0	163	470	17.53	31.99	0	193
GAMES	546	95.23	40.57	7	162	470	115.49	36.88	10	162

Econometric Techniques and Results

Given panel data, estimation of the model via OLS and pooled OLS may be inappropriate due to omitted variable bias, which occurs because immeasurable player characteristics in the

error term are potentially correlated with some regressors. OLS fails to utilize the fact that there are several observations on each player. Pooled OLS addresses yearly changes, but still ignores the cross-sectional player effects. Since the unobserved or unmeasured player effects α_i and v_i are accounted for in fixed effects (FE) and random effects (RE) models, omitted variable bias is addressed, but at the expense of reduced degrees of freedom relative to OLS or pooled OLS estimation. Both FE and RE models assume the unobserved player traits and the stochastic error terms are random variables. However, the FE models assume the player traits are correlated with some explanatory variables, whereas the RE model assumes no such correlation. FE models control for each player, which allow the within-player variation to be used to estimate the coefficients in (1). RE estimation also uses the between-player variation and the within-player variation in estimating (1).

There are three reasons why FE estimation of (1) is expected to be more suitable than RE estimation for this study. First, FE modeling assumes the unobserved or unmeasured effects are correlated with other regressors, whereas a RE does not. There are several potential areas for correlation. A player's ability, captured in α_i , is expected to be positively correlated with the number of games played, since higher ability players are thought to play in more games. Additionally, if a player has high ability, he will likely contribute more to his team success and increase his team's chances of making the playoffs, implying an expected positive correlation between α_i and PLAYOFF. If a player is retiring, it is expected it to be attributed to a family decision or health factor captured within v_i , implying suspected correlation between v_i and RETIRE, NOPLAY or PROBRETIRE. As discussed by Krautmann and Oppenheimer (2002) and Kahn (1993), one may expect negative correlation between being in a contract year, (CONTRACTYR) and unobserved risk tolerance within α_i or v_i because more risk-averse players will seek longer contracts, hence have fewer contract years between 2006 and 2011.

Second, Allison (2005) explains the standard errors for FE model estimates may be larger than those for RE models if there is little variation of predictors within individuals over time coupled with relatively greater variability across individuals. If so, higher p-values appear for coefficients in FE models than RE models, leading to potentially fewer statistically significant results in FE models. Perusing the data, one could see suitable variation within players' games played, team playoff chances, and OPS to suggest the FE approach's appropriateness. (More objective evidence of such is provided in the results section.) Lastly, ignoring the between-player variation in FE estimation eliminates the expected confounding effect of higher ability players having higher OPS statistics, thus allows for increased effort purported in CONTRACTYR, as opposed to ability, leading to an improvement in OPS. A Hausman test tests for the rejection of a RE model in favor of a FE model. It is expected that the Hausman test will show that FE is the most appropriate technique due to the three reasons above.

Table 3 provides the OLS and pooled OLS results for models using either OPS or OPS100 as the dependent variable. The first and second columns of each estimation technique

are NOPLAY, the ex-post retirement variable, and RETIRE, the inferred retirement status, respectively. The two-tailed p-values lie in parentheses below each coefficient. Following Krautmann and Solow (2009), the dummy variable DEFENSE = 1 captures hitters that play catcher or shortstop as they are relied upon for their defensive skills and not necessarily their offensive production. The OLS and Pooled OLS models show that catchers and shortstops have about a .033 to .038 lower OPS than the hitters playing other positions, ceteris paribus, compared to decreases in OPS100 between 9.4 and 10.5. These negative coefficients are in keeping with Krautmann and Solow (2009).

Most importantly, note the contract year coefficient is generally statistically significant and negative across all models. Since OLS estimation uses the variation between players and within players, the drops in the mean OPS and OPS100 for contract years highlighted in Table 2, lead to statistically significant negative coefficients for the contract year. The adjusted R^2 lies between .30 and .34 for the eight models. Although not shown, substituting years of experience for age led to similar results in all coefficients and adjusted R^2 's.

Table 3: OLS and Pooled OLS Results for OPS and OPS100 as Dependent Variables

DEPENDENT VARIABLE:	OPS				OPS100			
	OLS	OLS	POOLED OLS	POOLED OLS	OLS	OLS	POOLED OLS	POOLED OLS
N=1016								
INTERCEPT	1.076	1.082	0.9852	0.9629	159.6	161.777	180.016	175.036
AGE	-0.0287 (.0299)	-0.0295 (.0287)	-0.0233 (.0842)	-0.0223 (.1032)	-6.062 (.0710)	-6.285 (.0651)	-7.2928 (.0331)	-7.0828 (.0415)
AGE2	0.0004 (.0246)	0.0004 (.0282)	0.0004 (.0623)	0.0003 (.0853)	0.0937 (.0586)	0.0942 (.0611)	0.1095 (.0295)	0.1043 (.0410)
GAMES	0.0013 (.0001)	0.0014 (.0001)	0.0013 (.0001)	0.0014 (.0001)	0.317 (.0001)	0.3482 (.0001)	0.3173 (.0001)	0.3494 (.0001)
DEFENSE	-0.0376 (.0001)	-0.0333 (.0001)	-0.038 (.0001)	-0.034 (.0001)	-10.4909 (.0001)	-9.5167 (.0001)	-10.4113 (.0001)	-9.4381 (.0001)
PLAYOFF	0.0248 (.0002)	0.0281 (.0001)	0.025 (.0002)	0.0283 (.0001)	5.1317 (.0028)	5.8965 (.0007)	5.0778 (.0030)	5.878 (.0007)
NOPLAY	-0.0661 (.0001)		-0.0636 (.0001)		-15.142 (.0001)		-15.704 (.0001)	
RETIRE		-0.043 (.0058)		-0.0422 (.0067)		-10.7354 (.0064)		-10.8331 (.0060)
CONTRACTYR	-0.0138 (.0401)	-0.0197 (.0036)	-0.0129 (.0557)	-0.0181 (.0074)	-2.2242 (.1917)	-3.3578 (.0380)	-2.433 (.1538)	-3.712 (.0301)
YR			-0.0047 (.0416)	-0.0062 (.0078)			1.0606 (.0709)	0.6912 (.2420)
ADJUSTED R^2	0.3361	0.3149	0.3381	0.319	0.3173	0.3003	0.3188	0.3006

One-way fixed effect model estimation, FE1, controls for the time-invariant traits in a_i affecting OPS, whereas two-way fixed effect model estimation, FE2, controls for the time-variant traits in v_t as well. The player's identification number, CODE, controls for a_i while the year played within the six potential years of playing, YR, accounts for v_t . Both FE1 and FE2 were tested, but only the FE1 results are presented in Table 4 because none of the FE2 models showed significant coefficients on the YR's. As suspected, the Hausman test rejected RE modeling in all cases with $p < .0001$, suggesting only FE results apply.

To compare the FE1 estimates to the corresponding ones in Table 3, the first four columns of Table 4 pertain. Table 4 shows the R^2 's in are about two times greater than those in Table 3, which is not surprising since including a dummy variable for each player should increase the model's predictive power. The models using NOPLAY yield higher R-squares than those with RETIRE, consistent the Table 3 findings. The signs on the coefficients for PLAYOFF and GAMES are positive and statistically significant in Table 3 and Table 4. Using the first column of Table 4's GAMES coefficient of .0008 implies raising the OPS by .10, perhaps from the mean of .724 to .824, requires playing 125 additional games, which may be materially impractical. In Table 3, the corresponding coefficient on GAMES suggests 76 additional games needed. As expected, being on a playoff team increases the OPS and OPS100.

As in Table 3, the retirement status coefficients on NOPLAY and RETIRE in Table 4 are negative and statistically significant across all cases and the NOPLAY coefficients are always larger negative numbers than those on RETIRE. The FE1 estimation shows that for players who do not appear on MLB rosters the year following a contract year, their OPS is expected to be about .0661 lower, *ceteris paribus*, or 9.13% lower relative to the mean OPS. For the RETIRE-inclusive model, the estimated impact is -.0403 or -5.57% relative to the mean. For OPS100, the expected declines relative to the mean of 91.15 are 17.58% and 10.35% for NOPLAY and RETIRE, respectively.

There are glaring differences between the OLS and FE1 estimations. First, for OPS, the DEFENSE coefficients essentially flip signs from -.03 to .03, moving from OLS to FE1 estimation. Catchers and shortstops have a mean OPS of .68 compared to all others with a mean of .74, which is what the OLS and Pooled OLS estimations are capturing. The FE models appropriately control for the differences in unobserved characteristics for catchers and shortstops, and contrary to expectations, these players are expected to have a higher OPS, *ceteris paribus*. Interestingly, the p-values are larger for all the DEFENSE coefficients in the FE models compared to the OLS models. Further analysis reveals that 86% of the variation in DEFENSE is between-players and 14% within-players, leading to an increased the standard error for DEFENSE. As Allison (2005) notes, fixed effects is most efficient when all of the variation is within, not between, players. For PLAYOFF and GAMES the variations within-players are 75% and 47%, respectively, which are more sufficient for improved p-values with FE1 estimation.

Secondly, and most importantly, are the statistically significant positive coefficients on CONTRACTYR for the FE1 models with ex-post retirement, NOPLAY. For players in their contract year, the expected increases in OPS and OPS100 are .013 and 3.63, respectively, which represent 1.8% and 3.98% increases relative to their respective means. For the RETIRE-inclusive FE1 models, the contract year phenomena are smaller, an expected 1.2% improvement in OPS and 10.35% expected increase in OPS100. The analysis of variance table, not shown, associated with the first column of Table 4, indicates 55% of the variation in OPS is between-players and 45% within-players. Since 68.9% of the variation in OPS is attributable to the model and only 55% comes from across-players, the within-player attribution matters. The corresponding percentages for OPS100 in the fourth column are 59% and 41%. High percentages attributable to differences within-players reinforce the need to use FE1 estimation.

Table 4: One-Way Fixed Effects Models (FE1) for OPS and OPS100

DEPENDENT VARIABLE:	OPS	OPS	OPS100	OPS100	OPS	OPS100
INTERCEPT	1.929	1.834	270.644	245.83	2.749	496.692
AGE	-0.059 (.0083)	-0.0502 (.0307)	-10.2983 (.0701)	-8.0304 (.1730)	-0.1446 (.0001)	-33.1451 (.0001)
AGE2	0.0007 (.0411)	0.0005 (.1499)	0.1417 (.1013)	0.0945 (.2913)	0.0025 (.0001)	0.6291 (0.0001)
GAMES	0.0008 (.0001)	0.0009 (.0001)	0.1957 (.0001)	0.2356 (.0001)	0.0001 (.1718)	-0.1903 (.0001)
DEFENSE	0.0278 (.0701)	0.0306 (.0519)	7.7002 (.0488)	8.3647 (.0362)	0.0127 (.2228)	3.9655 (.1386)
PLAYOFF	0.0145 (.0197)	0.0178 (.0051)	2.6874 (.0891)	3.4829 (.0309)	0.0015 (.7289)	-0.2165 (.8419)
NOPLAY	-0.0661 (.0001)		-16.0262 (.0001)			
RETIRE		-0.0403 (.0079)		-9.4276 (.0144)		
PROBRET					-.6604 (.0001)	-168.26 (.0001)
CONTRACTYR	0.013 (.0354)	0.0087 (.1668)	3.6305 (.0213)	2.5742 (.1075)	0.0079 (.059)	2.2195 (.0382)
BUSE R ²	0.6892	0.674	0.6785	0.6641	0.8575	0.8492

In an effort to overcome the endogenous nature of ex-post retirement with OPS or OPS100, the probability of retirement, PROBRET, is estimated akin to Krautmann and Solow (2009). Using binary NOPLAY as the dependent variable precludes the use of a logistic regression model using FE estimation. Since NOPLAY is always the last observation for anyone "retiring" and YR increases each year until this last observation, it is not possible to find the

maximum likelihood estimate for a FE logistic regression. Essentially, any regressor that increases monotonically will predict the likelihood of retirement, leading to an invalid model fit with a singular matrix, as noted by Allison (2005). A way around this is to estimate a linear probability model using FE estimation.

The likelihood of retirement, i.e., not playing in MLB the following year, is expected to be positively related to the number of days on the disabled list in the current year and years of experience having played in MLB, while negatively correlated with OPS or OPS100. The model, not shown, using OPS (OPS100) yielding the greatest percent correctly predicted of .934 (.939) employs the quadratic form of years of experience with a FE2 estimation. The coefficients for days on the disabled list are positive, but not significant, in either model. The other variables show expected signs and p-values of .0001. For players whose probability of retirement lies below 0 or above 1, PROBRET is reset to .001 or .999, respectively. FE1 models using the predicted values of retirement, PROBRET, comprise the last two columns of Table 4.

If greater offensive performance reduces the likelihood of retirement, which in turn is associated with more effort and greater offensive performance, one expects a positive bias for the NOPLAY and RETIRE coefficients. Mitigating the bias by creating PROBRET should lead to PROBRET coefficients falling in absolute value terms, as shown in the last two columns of Table 4. A one percentage point increase in the likelihood of retirement is expected to decrease OPS and OPS100 by .0066 and 1.6826, respectively, *ceteris paribus*. As percentages of their means, the declines are .912% and 1.85%. While DEFENSE and PLAYOFF lose their statistical significance in both models, most of remaining predictors show expected signs and greater statistical significance compared to the first four columns. The contract year phenomenon remains highly significant with 1.09% and 2.43% expected increases in OPS and OPS100, respectively. Lastly, models with PROBRET boost the Buse R^2 by nearly .20.

Conclusions

By comparing regression models using OLS and Pooled OLS versus Fixed Effects, only the latter yields evidence of the contract year phenomenon. Given multiple observations per player, this information is appropriately utilized using FE estimation by controlling for unobserved traits (e.g., ability, risk tolerance, familial conditions, etc.) within players that are expected to be correlated with observed predictors (e.g., games played, playoff contention, etc). OLS techniques cannot isolate the within player impacts of the regressors on offensive performance. The OLS and Pooled OLS techniques capture the surprising decline in the mean OPS for contract years and inappropriately suggest a negative contract year impact.

Using FE estimation, the two hypotheses regarding contract year performance are confirmed. First, the results lead one to believe players increase their effort to boost their performance in their contract year so as to garner another desired contract. The OPS of a free

agent hitter in his contract year is expected to be 1.09% to 1.8% greater than his OPS in a non-contract year, depending on the form of the retirement variable that is held constant. O'Neill and Hummel (2011) using data from 2004-2008 find larger contract year impacts on OPS. For OPS100, the contract year boost estimates lie between 2.43% and 10.35%. Using the Grossman et al 2004 heuristic that every .100 increase in OPS raises salary by \$2,000,000 implies the contract year boost is expected to increase annual salary between \$158,000 and \$260,000, which represent 5.1% to 8.4% of the average 2011 MLB average salary. While still materially impactful, the larger contract year boosts noted by O'Neill and Hummel (2011) suggested greater pay checks. Perhaps over the 2006-2011 CBA period team owners sent a signal that contract year boosts are not valued as much as in the past as predicted by Dinerstein (2007).

"Retiring" players do not have a financial incentive to increase their effort, although they may have a desire to go out on top. Accounting for retirement is necessary as a predictor for OPS since the lower mean OPS for the contract year is due to lower OPS values for players no longer on a MLB roster the following year. Models ignoring retirement will be misspecified. The second hypothesis-confirming finding from the FE1 results indicate that a player who "retires" after the contract year season is expected to have between a .91% and 9.13% decline in his OPS, ceteris paribus. For OPS100, the retirement effect decreases range from 1.85% and 17.58%. The smallest retirement impacts occur when employing the bias-mitigating, predicted likelihood of retirement variable. The increased goodness-of-fit and bias reduction associated with PROBRET models suggest PROBRET is the preferred retirement measurement.

For future research, the flip side of contract year opportunistic behavior, shirking, will be examined. Using the same players, the shirking hypothesis - players who have secured a long term guaranteed contract will reduce their effort and hence performance after signing the contract - will be tested. It is of interest to see if the same players who engage in the contract year performance boost are the same as those who shirk with a new contract.

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