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Initial Public Offerings of Ballplayers

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Abstract

As a field study of choice under uncertainty, we examine baseball teams' investments in amateur players. Though most prospects fail to deliver any return on their multi-million dollar signing bonuses, returns on the minority who succeed easily offset these losses: the expected annual yield on the median first-round draftee is 33 percent. However, the pattern of returns is inconsistent with market efficiency. Yields are lower for high schoolers than collegians (27 percent vs. 43 percent), lower for pitchers than position players (24 percent vs. 41 percent), decline for later round long-shots, and may be negative under competitive bidding.

JEL Classification Codes: D8, G14

Keywords: Market efficiency; Bounded rationality; Prospect theory; Winner's curse

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"If professional baseball players, whose achievements are endlessly watched, discussed and analyzed by tens of millions of people, can be radically mis-valued, who can't be? If such a putatively meritocratic culture as professional baseball can be so sloppy and inefficient, what can't be?"

--Michael Lewis, author of *Moneyball: The Art of Winning an Unfair Game,* as quoted in Neyer (2003)

I. Introduction

Each year, major league baseball teams risk enormous sums to sign amateur players to their first professional contracts. For example, 29 "blue-chip" prospects selected in the first round of baseball's 2004 amateur draft received guaranteed signing bonuses totaling nearly $57 million. In the great majority of cases, however, teams receive little or no return on these investments. According to one reputable survey (Callis, 2003), just one of every four first-round picks ultimately makes a non-trivial contribution to a major league team, and a mere one in twenty becomes a star. Though signing bonuses are lower for players drafted in later rounds, so are the chances that these investments will pay off. All told, only eight percent of players drafted in the first ten rounds (of a 50-round, 1,500-player draft) typically develop into big league regulars, and only after a lengthy apprenticeship in the minor leagues. In addition, the sport's annual draft, which clearly aims to limit players' bargaining power by granting their drafting teams exclusive negotiating rights for a year, only applies to amateur players in the United States. Foreign-born players—now about a quarter of those on major league rosters and about a third of those in the minors—may bargain with any team prior to signing their initial professional contracts, and thus often negotiate bonuses equal to or greater than those of top domestic draftees (though data on the amounts paid to such players and their chances for eventual success are sketchy).

This high-uncertainty, high-stakes market provides a unique opportunity to examine the way individuals value risky choices. There is abundant experimental evidence that when people form judgments about probability they depart systematically from the laws of statistics (see Camerer, 1995, for a review of this literature and Camerer, 1998, for an update). There is also experimental evidence that—in auctions involving considerably less uncertainty than in the baseball labor market—bidders are subject to a "winner's curse" in which they realize below-normal profits because they systematically over-value available goods (see Kagel, 1995, for a review of this literature, and Cox, Dinkin, and Smith, 1999, for an argument that this evidence is more limited than previously supposed). Critics of the experimental approach argue, however, that the incentives commonly present in a laboratory setting may be inadequate to induce subjects to incur the costs that optimal behavior might require, and that the real world provides both greater incentives to behave efficiently and more opportunities to learn how. Thus, there is much to be learned from field studies of economic decision-making. It is in this spirit that we offer the present study.

There is certainly reason to suspect that behavior in the market for aspiring professional ballplayers might be consistent with the theory of bounded rationality, which stresses that agents often have incomplete knowledge of available alternatives and computational abilities too limited to solve for theoretical optimum values (Simon, 1986), and/or with prospect theory, one implication of which is that low-probability events tend to be overweighted by decision-makers (Tversky and Kahneman, 1986). Decisions about whom to draft or sign domestically and internationally are based on information from experts—“scouts”—who directly observe only a fraction of the available pool of talent, and often for just a few games. There is anecdotal evidence that scouts bring certain biases to the evaluation process and are excessively influenced by their most recent observations of their subjects.[1] Further, in conveying their evaluations of available players to the ultimate decision-makers, scouts often use analogies linking prospects' traits to those of established major-leaguers. While this practice facilitates economical communication, it might contribute to overestimation of an individual amateur's real probability of making the majors; indeed, given the data summarized earlier, one wonders why clichés like “blue-chip prospect” or “can't-miss prospect” are even part of baseball lexicon. Finally, teams face an information asymmetry in judging the health, motivation, and, remarkably, even the age[2] of a prospect.

Accordingly, our primary task in this study is to assess whether the bonuses teams pay for available amateur talent are efficiently priced. In the next section, we begin by summarizing relevant characteristics of the baseball labor market. In brief, the sport's collective bargaining agreement grants teams a fairly lengthy period over which they may recover the costs of acquiring and developing talent by paying players wages below their net marginal revenue products. Section III details the methodology and data used to

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[1] For an “inside” look at the scouting and drafting functions, see Lewis (2003), chapters 2 and 5. A best-selling business journalist, Lewis was granted unprecedented access to the front office of the Oakland Athletics, a team committed to discovering and exploiting market inefficiencies in order to compete with better-financed rivals.

[2] Amid heightened security checks following the 9/11 terror attacks, it was found that numerous foreign-born players had understated their true ages when signing their initial contracts. The incentive to do so is clear: scouts forecast a player's future potential by making age-adjusted assessments of ability, so bonus payments are generally higher for younger players of a given perceived talent level. In 2002, for example, the Arizona Diamondbacks were willing to pay $500,000 to sign a pitcher they thought was a 17-year-old named Adriano Rosario of the Dominican Republic. In 2004, they learned that the player was actually three years older and named Tony Pena; he had assumed the identity of a younger nephew at the time he was signed. See Price (2004).

examine the benefits and costs of teams' investments in amateur players, and Section IV presents results. We find that though three-quarters of first-round draftees yield zero returns, the rest generate payoffs high enough to justify the enormous bonuses paid. On average, the expected annual return on these investments approaches 33 percent.

Given the monopsony power inherent in the draft system, however, finding that yields on top picks are attractive probably should not be a surprise and certainly is not a powerful test of the rationality of bidders; rather, it is a benchmark useful for evaluating their other choices. What is surprising, and suggestive of irrationality, is that we find differences in expected returns for players from various cohorts, and that these differences are contrary to the predictions of financial theory. Teams pay significantly higher bonuses to high school first-round draftees despite the fact that college selections deliver returns more quickly. As a result we calculate an IRR of 43 percent for college selections and 27 percent on high school draftees. In addition, the yield on pitchers is 24 percent and 41 percent for position players—despite the fact that the former face considerably greater injury risk than the latter. Finally, expected returns fall precipitously for the lower-quality investments that teams make in their second- and third-round draftees. In sum, it appears that teams systematically over-estimate many prospects' probability of success and/or fail to appropriately adjust their bets to take account of longer odds.

In Section V, we present some anecdotal and historical information about bidding behavior outside the context of the draft (i.e., under more competitive conditions). We find that when teams lack the monopsony power the draft confers, it is not at all unusual to find that bonuses rise quickly to levels that drive expected returns below zero. Section

VI summarizes our findings—which are quite consistent with the experimental literature supporting bounded rationality or prospect theory[3]—and offers concluding remarks.

II. Institutional background and basic model

When a major league baseball team pays a multi-million dollar signing bonus to a top draft choice it is placing a large bet, returns on which are both highly uncertain and long deferred. Even if a draftee beats the odds (which are, again, 3-to-1 against a first round pick, lengthening to 9-to-1 against a second rounder and 15-to-1 against a third-rounder) and becomes a contributing big-league regular, this accomplishment generally will follow a minor-league apprenticeship that almost always lasts years. Should this apprenticeship produce a successful major leaguer, however, his employer will earn sizeable monopsony rents thanks to the collective bargaining agreement that governs the baseball labor market. Competitive bidding may drive wages to equality with workers' marginal revenue products in other labor markets, but ballplayers cannot auction their services to the highest bidder until they have accumulated six years' major league service. During this monopsony phase of a player's career, he usually will be paid far less than his marginal revenue product.

Prior studies (e.g., Marburger (1994, 2004) and Burger and Walters (2005)) have estimated that employers keep over two-thirds of a player's marginal revenue product for

[3] Our findings are also consistent with empirical literature that has found over-bidding on initial public offerings (IPOs) of shares by previously closely-held firms. For example, Ritter (1991) found that IPO issues during 1975-84 substantially underperformed a sample of matching firms from the closing price on the first day of trading to their three-year anniversaries. Ritter and Welch (2002) reached a similar conclusion for issues during 1980-2001. In that sample, over-eager buyers drove up first-day prices by an average of 18.8 percent, so that an investor buying shares at the first-day closing price and holding them for three years underperformed the market by 23.4 percent. Evaluating the copious literature seeking to explain this market anomaly is well beyond the scope of this paper, but our findings hint that cognitive psychology might have a part to play.

those with three or fewer years of big-league service, with the exploitation rate declining thereafter as a result of an arbitration process that is also a by-product of baseball's collective bargaining agreement. Only after this six-year period may a "free agent" player auction his services to the highest bidder, presumably receiving rewards commensurate with his expected net marginal revenue product.[4] The primary question of this study is whether the rents extracted from players prior to their eligibility for free-agency are high enough to justify both the initial signing bonuses (paid not just to those who become big-league regulars but to the majority of draftees who never make the majors) and the lengthy delay in receipt of these rents.

As illustrative examples of the rewards and risks of the baseball draft, consider the first players selected in 1990 and 1991. Larry “Chipper” Jones received what appears in retrospect to be a bargain-basement signing bonus of $328,000 (in real, 1998 dollars) from the Atlanta Braves in 1990. Jones's minor-league apprenticeship (which included some down time due to injury, another significant element of risk in this market) lasted until 1995, when he became a big-league regular. According to reasonable estimates of his output and the value of that output to his employer (explained in more detail below), Jones's marginal revenue product for the Braves was $6.13 million in 1995 (again, in real, 1998 dollars) and averaged $8.4 million for the next five seasons. But his real salary increased from a mere $120,000 in 1995 to $4.59 million by 2000. Thus, thanks to the unique structure of the baseball labor market, the Braves realized significant rents during each of Jones’s first six big-league seasons, summarized in Figure 1. By contrast, in

[4] Exactly why the players' union has ceded such monopsony power to teams in successive collective bargaining agreements is a subject of some debate. While some argue that it is a way to allow teams to recover presumed costs of developing players during their apprenticeships, the union has also expressed the belief that limiting the availability of free agency raises overall wages, at least to star players. See Fort (2003, p. 259).

1991 the Yankees paid Brien Taylor a then-record bonus of $1.8 million (in 1998 dollars). But due to an off-field injury to his pitching shoulder two years later, Taylor never appeared in a major league game, a fate shared by over one-third of top draft picks and two-thirds of those drafted in the first ten rounds. The Yankees' enormous investment in Taylor was a total loss.

< Insert Figure 1 about here >

In what follows, we evaluate teams' draft investments by calculating an internal rate of return (IRR) on their bonus outlays. Rather than calculate returns on specific players like Jones or Taylor, however,[5] we employ historical data to make generalizations about the probability that a particular draft pick will pay off, the normal waiting period for doing so, and the typical size of realized returns. In general terms, the payoff for a successful draft choice will be a stream of future rents *R* per year *t*, where:

$$R_t = NMRP_t - w_t \,, \qquad \textbf{(1)}$$

and *NMRP* denotes a player's net marginal revenue product and *w* his wage.[6] The present value (*PV)* of these receipts during a player's monopsony period is then simply:

$$PV = \sum_{t=t_0}^{t_0+5} \left[\frac{R_t}{(1+i)^t} \right], \qquad \textbf{(2)}$$

where t_0 represents the interval between the date the player is drafted and the time it takes him to reach the majors and *i* denotes the appropriate discount rate.

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[5] Calculating individual returns is problematic for several reasons, chiefly (a) unavailability of information on individual bonuses for sufficiently many players to produce an adequate sample size, and (b) the computational issues arising when such a large fraction of investments produce an IRR of -100 percent. Thus, we consider a probabilitistic approach to be superior in these circumstances.

[6] While it is true that there are additional player development costs during a prospect's apprenticeship period, these are not relevant to our evaluation. The costs of coaching and the modest salaries paid to minor leaguers, for example, are the same for a top draft pick as for a 50th-round pick; the question before us is whether it makes economic sense to pay a large bonus to secure a "blue chip" prospect, for minor league rosters can be filled out with lower-round picks not requiring such payments.

Since a draftee may never make the majors, however, and since the quality of his performance may vary considerably even if he does, equation (2) must be refined. The probability- and quality-weighted present value of the expected rents generated by a draft choice may be expressed as:

$$PV = \sum_{j} p_j \sum_{t=t_0}^{t_0+5} \left[\frac{R_{t,j}}{(1+i)^t} \right], \qquad \textbf{(3)}$$

where p_j is the probability of drafting a player of quality j.

In sum, we can calculate the IRR on draft-choice investments if we can estimate the probability (p) that draftees will make the majors, the length of time it typically takes to do so (t_0), and the magnitude of the annual rents (R) successful draftees deliver to their teams. The latter will, of course, depend on the quality of their performance (j) over the six years prior to eligibility for free agency and the rate of monopsony exploitation during this period. Setting PV equal to the typical bonus outlay, we can solve for the discount rate i (the IRR) on such bonuses.

III. Detailed methodology and data

Applying this approach to a statistically meaningful sample of players requires that, first, we quantify the probability that a given draft choice will become a productive major league player and, second, we differentiate drafted players according to their ultimate quality. Fortunately, Callis (2003) has done both, evaluating 2,115 players drafted in the first ten rounds over eight years (from 1990 to 1997). Callis, a veteran talent evaluator and executive editor for a leading baseball journal, identified those who failed to make the major leagues (67.1 percent of his total sample, which he dubbed "flops") and sorted successful draftees into five qualitative categories, from those who merely "got a cup of coffee" in the majors (17.1 percent), to "fringe" players (7.8 percent), on up to “stars.” For our purposes, only the top 3 categories are relevant: those who became "regular" (5.1 percent of the total sample), "good" (2.1 percent), or "star" quality players (0.9 percent). In general, draftees who fail to become at least major league regulars will not generate positive cash flow for their teams.

Given the Callis (2003) sample of 1990-1997 draft choices, we then are able to calculate the number of years between draft day and major league debut (t_0) for players in each of the three relevant quality classifications. Since signing bonus information for later rounds is less available and less accurate, we focus on the first three rounds of the draft. As we might expect, stars tend to reach the majors sooner than good players, who in turn usually serve shorter apprenticeships than regulars. Table 1 displays this and other summary statistics for first-round draftees; later, we will discuss the characteristics of second- and third-round draftees and examine other issues teams face in making their selections and bidding appropriately.

We next quantify the contributions of players of the relevant quality levels to their teams' performance, a necessary step toward calculating their marginal revenue products and contributions of rents to their teams. In our view, the best overall measure of players' physical productivity is Bill James' Win Shares statistic (James and Henzler, 2002), which aggregates all of the myriad ways a player can contribute to his team's performance (with his bat, glove, arm, and legs) into a single statistic, which is effectively the player's marginal contribution of wins to his team. This measure also has the virtue of dovetailing neatly with the traditional economic approach to measuring players' marginal revenue products, devised by Gerald Scully (1989), which involves multiplying a player's marginal output of wins by his team's marginal revenue from extra wins. We tabulate the marginal output of wins by each successful draftee in the three relevant quality categories; as Table 1 shows, on average stars were nearly twice as productive as good players, who were in turn twice as productive as regulars.

< Insert Table 1 about here >

We next combine these estimates of players' marginal output of wins with estimates of the marginal value of wins to their teams in order to quantify players' marginal revenue products during their first six years in the major leagues. This requires estimation of a revenue function for a representative team.[7] We use data on local revenue over 1995-2001 for all major leagues teams, available from the Commissioner's Blue Ribbon Panel on Baseball Economics (Major League Baseball, 2000) and a later update (Major League Baseball, 2001), to estimate the following linear revenue function via ordinary least-squares regression:

$$\underset{}{TR} = \underset{(17.79)}{-92.02} + \underset{(1.25)}{6.18t} + \underset{(1.75)}{6.96M} + \underset{(0.21)}{1.15W} + \underset{(9.33)}{38.41STAD} - \underset{(1.53)}{2.63AGE} + \varepsilon. \qquad \textbf{(4)}$$

Standard errors are in parentheses, while *TR* signifies local team revenue in millions of inflation-adjusted dollars (base year = 1998), *t* is a time trend, *M* is a measure of market size (SMSA population, in millions), *W* signifies team wins, *STAD* is a dummy variable which takes a value of one if the team plays in a new, state-of-the-art stadium, and *AGE* is the number of years such a stadium has been in use. The coefficient on *W* implies the

[7] Burger and Walters (2003) showed that both the intercepts and slopes of teams' revenue functions vary according to the size of their markets, so marginal win values are higher in larger markets. At this point, however, since our goal is to calculate IRRs for representative players taken in different draft rounds, we rely on an "average" marginal win value. We will return to the question of how market size affects bidding behavior later, in Section 4.

marginal value of a win was $1.15 million on average during the 1995-2001 sample period.

Based on this estimate, we could calculate players' gross marginal revenue products, á la Scully, simply by multiplying their marginal output of wins by this $1.15 million marginal-revenue-per-win figure. However, Burger and Walters (2005) find that free agents' market-determined salaries are typically about 80 percent of their gross marginal revenue products. In effect, teams appear to adjust downward the wages they are willing to pay in competitive markets in recognition of non-trivial marginal costs of employing particular players—most significantly, perhaps, the costs of insuring (or self-insuring) against injury. In this study, failure to take this fact into account might produce upward-biased estimates of the rents that may be extracted from successful draftees, and upward-biased estimates of the rate of return on investments in draftees in general. Accordingly, we estimate the net marginal revenue product (*NMRP*) of successful draftees as:

$$NMRP_t = MP_t * (0.8 * MWV)\,, \qquad \textbf{(5)}$$

where *MP* denotes the player's marginal product in wins, and *MWV* equals the marginal revenue delivered by each extra win for a representative team.

Finally, to calculate estimated wages for successful draftees, we first assume these players earned the league minimum salary in years one and two of their careers (when they are not eligible for arbitration, and therefore face their teams' unrestrained monopsony power). Then, in years three through six of players' careers, we estimate that their wages rise (and their monopsony exploitation rate declined) in accord with the typical ratio of the salary awarded by arbitrators to a player's net marginal revenue product, as presented in Burger and Walters (2005). The equation for a player's estimated wage (w) in each year of the monopsony phase of his career is thus:

$$w_t = \begin{cases} \min \ldots\ldots\ldots\ldots\ldots\ldots t = t_0; t_0 + 1 \\ 0.375 * NMRP_t \ldots\ldots t = t_0 + 2 \\ 0.550 * NMRP_t \ldots\ldots t = t_0 + 3 \\ 0.763 * NMRP_t \ldots\ldots t = t_0 + 4 \\ 0.800 * NMRP_t \ldots\ldots t = t_0 + 5. \end{cases} \qquad \textbf{(6)}$$

Based on the foregoing data and assumptions, we proceeded to evaluate the efficiency of teams' decisions to invest in high draft picks during the late 1990s.

IV. Results

Table 2 presents the typical stream of rents that a representative team can expect to extract from successful draftees in the three relevant quality categories. For example, a star-caliber player may be expected to produce a net marginal revenue product (*NMRP*) of $7.4 million per year (or 80 percent of his expected gross marginal revenue product, equal to roughly eight marginal wins per year times $1.15 million per win) during the monopsony phase of his career. This phase normally begins 2.22 years after he is drafted, after which his estimated wage (w) rises from the league minimum to roughly $3.79 million, yielding rents (R) to his team ranging from $7.1 million down to $3.6 million per year.

< Insert Table 2 about here >

Thus, drafting a future star delivers enormous returns—but since the probability that even a first round draft pick will become a star is a mere 4.3 percent, and since both future stars and flops will command multi-million dollar signing bonuses, evaluating the wisdom of such investments requires the calculation of probability- and quality-weighted present values of these prospective rents (as specified by equation (3), above). Our preferred method of evaluation involves setting the present value expression (equation 3) equal to the typical bonus payment and solving for the internal rate of return (IRR) on draft investments.

In order to perform a benchmark IRR calculation we must settle on an appropriate benchmark first-round bonus. A high degree of variation in team revenue over time and in bonuses paid by draft rank within rounds represent complicating factors. Our marginal win value of $1.15 million is based on equation (4) estimates of the revenue function using 1995-2001 data. For consistency, we base our IRR calculations on the median bonus for this same time period. First, we model the variation in first round bonuses within and across years (1995-2001) by estimating the following regression[8]:

$$\ln(realbonus) = \alpha + \beta_1 Year + \beta_2 Slot + \beta_3 Slot^2 + \varepsilon\,, \qquad \textbf{(7)}$$

where bonus data are in real 1998 dollars, *Year* allows for a time trend, and *Slot* specifies the draft position (one for the first pick, two for the next, and so on). We model the natural log of bonus payments because a simple linear regression suffers from serially correlated errors. The square term on draft slot allows for the observed nonlinear decline in draft bonuses as one progresses deeper into the first round.

Coefficient estimates for equation (7) are presented in Column (1) of Table 3. This simple model explains over 68 percent of the variation in first-round bonuses for our seven-year sample. The *Year* coefficient indicates an annual rate of increase in real bonuses of almost 12 percent, holding draft position constant. The *Slot* coefficients reveal a nonlinear decline in bonus payments based on draft position. Taken together, the

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[8] Following Callis (2003), we define the first round to include supplemental picks. Unfortunately, bonus data were not available for supplemental first rounders in 1995-97, but were available for 1998-2001.

coefficients suggest that dropping one slot costs a player 5.5 percent at the top of the first round, but only 3.5 percent in the middle of the round (at the 20^{th} slot).

< Insert Table 3 about here >

A benchmark IRR: The median first-round pick

The midpoint of our bonus sample is 1998, and the midpoint of round one (including supplemental picks) in 1998 is represented by pick 22. Combining these values with the coefficient estimates of equation (7) yields an estimate of $1.02 million for the first-round median bonus payment. Setting *PV* in equation (3) equal to that figure, we find that the IRR (*i*) for the representative first-round pick is just under 33 percent.

Despite the fact that over three-quarters of first-round choices are failures, then, we find that the successes generate rents large enough to amply repay the lavish signing bonuses paid to all. Note, however, that this is only weak evidence of rational bidding by teams. While the estimated annual rate of return on the typical first-round bonus seems attractive, one must remember that (a) it is certainly enhanced by the monopsony power granted to teams by virtue of the draft's structure and (b) the high risk and unusual distribution of returns in this market prevent a definitive judgment about whether this benchmark IRR is comparable to alternative available investments. All we can say about behavior in this unique market thus far is that teams are both willing and able to assess their risks and constrain their bids in such a way that the dollars they invest in first-round draftees generate a yield of roughly 33 percent.

But history suggests important differences in the characteristics of draftees from different cohorts. For example, players with college experience are more physically mature, and their future performance may therefore be easier to forecast; players at certain positions also face greater risk of injury. The key question is whether teams are able to make appropriate choices or bids given such variety in the assets available to them. We proceed by testing for heterogeneity in characteristics of players from different cohorts, and when statistically significant differences are found (e.g., in the probability of success) we replace the benchmark average characteristics and re-compute the IRR. Our results show that returns on certain types of players are far higher than on others, hinting at the existence of market inefficiencies and suggesting that teams could improve their performance by pursuing undervalued assets or avoiding overvalued ones.

The high school anomaly

The benchmark IRR of 33 percent for first-round draftees is an average return on players from two separate talent pools with distinctive characteristics. At the least, high school draftees are younger and less experienced than those who have chosen to attend college rather than sign professional contracts. Some organizations favor drafting the former, thus enabling them to begin training their prospects at a relatively young age, but this strategy has been questioned by analysts who believe that focusing on college players mitigates information problems (leading to fewer drafting errors) and reduces the length of apprenticeships in the minors.[9]

< Insert Table 4 about here >

Table 4 displays summary statistics for first-round draft choices from high school and college. The probability of drafting a "regular" or better is very similar at 27.6

[9] Bill James, one of the founding fathers of the field known as "sabermetrics"—the statistical analysis of baseball—has famously opined that "college players are a better investment than high school players by a huge, huge, laughably huge margin" (as quoted in Lewis, 2003, p. 99).

percent for high school draftees and 26.8 percent from the college pool. Employing Z-tests for the difference between two proportions we fail to reject the null hypothesis of equality in probabilities within each quality class. The only significant difference in Table 4 is displayed in the middle column: college draftees reach the majors more quickly, as one might expect.

Given the greater speed with which college players make the majors, one might expect that teams would be willing to pay a premium to sign a player with college experience; our bonus regression, equation (7), allows a test of this hypothesis. The second column of Table 3 reports, instead, a negative and statistically significant coefficient on the college dummy variable, implying that teams selecting a college player receive roughly a 10 percent discount, holding draft position constant. The model predicts a median bonus of $1.068 million for a high school player and $961 thousand for the median college player.[10] After factoring in this college bonus discount and the shorter deferment period we use equation (3) to calculate a 43 percent IRR for college players and a 27 percent return for high school selections.

This finding of lower returns on high school selections lends some credibility to the analysts' critique that teams tend to overvalue high school players. That is, they appear to spend too many of their first-round picks on high school players, and/or they fail to adjust downward the bonuses they are willing to pay high school draftees to compensate them for the longer wait for a payoff.

[10] Higher average bonuses for high school players could result from a greater propensity to draft them earlier in the first round, but we find no systematic differences in the average draft positions of high school and college players from 1995-2001. It is likely that the higher price of high school draftees (holding slot position constant) reflects their enhanced bargaining power—i.e., the fact that they can threaten to attend college if offered bonuses are deemed inadequate. In any case, teams should rationally expect such threats and select from the two talent pools until anticipated returns from each are comparable.

In fairness to the teams, however, we should note that time series evidence on drafting behavior shows evidence of learning regarding the higher expected returns on college draftees. In 1971, for example, *every* first-round draftee was a high school player, but by 1981 the majority of first-round picks were collegians for the first time. Over the 1990-'97 sample period used for this study, the proportions from each pool were roughly equal, though recent years have seen a greater predilection for college players. In the 2004 draft, for example, 17 of the top 30 picks were collegians, and in 2005 20 of the top 30 picks were collegians. It is too soon to tell, of course, whether expected returns on the two cohorts have achieved expected equality, but there is reason to suspect they are moving in that direction.

The pitching anomaly

Choosing between the college or high school talent pools is not the only complex decision a drafting team must make. There are also important distinctions between players of different positions, especially between pitchers and regular position players. Pitchers suffer more frequent and more severe injuries than other players, suggesting that spending a high draft pick on a pitcher is a particularly risky proposition. Slicing the Callis (2003) draft data another way enables us to quantify the likelihood of successful pitching draft choices vis-a-vis other draftees. Table 5 indicates that the probability of drafting a player of good quality in the first round is significantly higher for position players relative to pitchers.

< Insert Table 5 about here >

Adding a pitcher dummy variable to our bonus regression, in column (3) of Table (3), we find that there is no evidence that teams reduce their bonus offers to drafted pitchers to offset the lower probability of success.[11] Using the median first-round bonus of $1.02 million and the significant probability differences from Table 5, we find that the IRR for pitchers drafted in the first round is a mere 24 percent, versus 41 percent for position players.[12] We conclude that teams over-sample and over-pay the population of pitchers available to them in the first round, irrationally foregoing higher-yield investments in position players.

The later-round anomaly

As the draft proceeds, teams draw from a dwindling supply of talent. Table 6 presents bonus information and success probabilities for first-, second-, and third-round draft choices. Clearly, there is a precipitous drop in the probability of drafting a good player in the lower rounds—but bonus payments appear to drop less dramatically. Indeed, using equation (3) and appropriate data to solve for IRRs by draft round, we find that returns drop from 33 percent for first-round picks to 20 percent for second-round picks to 13 percent for third-round picks. The declining observed yields suggest that teams hold irrational expectations about these lower-quality prospects' chances for future success, and/or that they fail to appropriately reduce the prices they are willing to pay for other, unknown reasons.[13]

[11] Nor is there any evidence that pitchers tend to be drafted later in the round than position players.
[12] These calculations assume the time to reach the majors and physical productivity within each quality class are equal for pitchers and non-pitchers. Productivity of pitchers from the Callis data is actually *lower* than regulars---which would exacerbate the anomaly---but the number of observations is very limited.
[13] By using the median bonus for each round we reduce the impact of outlier bonuses paid to the occasional first-round-caliber player who falls to lower rounds because of "signability" concerns by teams.

< Insert Table 6 about here >

A possible market-size anomaly

Prior research by Burger and Walters (2003) and Krautmann (forthcoming) has shown that a player's marginal revenue product will be positively—and quantitatively significantly—affected by the size of the market in which he plays. In the context of the present study, this means that rents earned on successful draftees will be far greater in large markets than in small ones. It would seem rational, therefore, for smaller-market teams to negotiate more aggressively to secure lower bonuses in order to achieve IRRs commensurate with our estimated benchmark, which presumably reflects an industry-average opportunity cost of capital.

Introducing a market size variable into our draft bonus regression, equation (7), allows us to test for this relationship. The results, reported in column (4) of Table 3, indicate that there is no evidence that market size affects the size of bonuses that teams are willing to pay their first-round draftees. It appears, then, that in this market bidders pay similar prices for assets which have significantly different values (given differences in their revenue-producing capacity). This might be understandable if draft picks were tradable, since teams which attach low value to draftees could simply swap them to teams where they have greater value—but for all practical purposes major league rules prohibit such transactions. The absence of a "market-size discount" therefore appears to be another anomaly of this market.[14]

[14] There is speculation that small-market teams avoid drafting players who have specified a high pre-draft price, instead choosing to draft cheaper players of lower perceived quality. If true, there would be market-size-related differences in teams' draft success rates. We found, however, a zero correlation between the

V. Anecdotal and historical evidence on bonuses under competition

As we have noted earlier, the structure of baseball's annual draft[15] grants teams significant monopsony power, enabling them to keep draftees' bonuses lower than they might be if there was open, competitive bidding for available players. Recent draft history provides illustrative evidence on this point. In 1996, four first-round draft picks took advantage of a loophole in the draft's rules, becoming free agents because their drafting teams had missed a specified deadline for making contract offers.[16] In the frantic bidding that resulted, the four players received bonuses that *averaged* over $5 million, more than two-and-one-half times the amount paid to the first (and otherwise-highest-paid) selection in that year's draft. What is more important is that this average investment far exceeded the "break-even bonus" of $3.25 million that, based on our calculations, would produce an expected IRR of zero for the typical 1998 first-round pick (given the relevant probability of making the majors, likely length of minor league apprenticeship, etc.). Unless there was genuine reason to fix these players' chances of success far above those of the typical first-rounder—and in hindsight none have become stars and only one has become a good major league regular—then their bonuses

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market size of teams and their success in drafting star, good, and regular players over the 1990-'97 draft years covered by the Callis (2003) sample.

[15] The drafting team has sole negotiating rights to the player and must offer him a contract within 15 days of selection. Failure to do so ends the club's rights and the player becomes a free agent, eligible to negotiate with any team. If the player attends a four-year college, the club's negotiating rights are lost as soon as the player attends his first class at the end of the summer, but no other club can sign the player until and unless it selects him in a succeeding draft.

[16] The players' drafting teams had missed the aforementioned (note 16) 15-day deadline for tendering contract offers. The players (and the bonuses they eventually received, in nominal dollars) were second pick Travis Lee ($10 million), fifth pick John Patterson ($6.075 million), seventh pick Matt White ($10.2 million), and twelfth pick Bobby Seay ($3 million). Source: Simpson (2005).

rationally would have been expected to generate substantial negative returns. In other words, the winning bidders in these four auctions were cursed.[17]

Of course, four observations do not a robust test make, but there are many other examples of apparently-extravagant bids for international players, to whom the draft does not yet apply. Several players have been paid amounts far in excess of the break-even bonus for first-rounders shortly after defecting from Cuba, which has well-developed amateur baseball leagues.[18] Similarly-large bonuses are routinely paid to amateur players from other Latin American countries, and recently Asian players have auctioned their services to major league teams for sizeable sums, as well.[19] The problem for researchers is that bonuses paid in these cases are not reported systematically. Teams guard the information for competitive reasons, and journalists generally take an interest only in players who succeed in making the major leagues, or nearly do so. There is insufficient data on the population of international signees who fail to deliver a return on teams' investments in them, so it is impossible to measure success probabilities and average apprenticeship periods in order calculate IRRs for this cohort. Nonetheless, the fact that

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[17] Another interesting winner's curse-type question arises from the fact that teams sometimes forfeit top draft picks when they choose to sign veteran free agents. Since bidding for such free agents is competitive, one might expect that the returns on such investments are, at the least, bid down toward "normal" levels, and certainly below the 33 percent IRR we have estimated for the median first-rounder. Accordingly, forfeiting such picks should be rare—unless free agent salaries are depressed enough to equalize IRRs across the two markets. We suggest this as a topic for future research: since some free agents can be signed without forfeiting draft picks while others cannot (depending on several arcane rules of baseball's free agency system), it would be interesting to see if prices for the latter are appropriately discounted.

[18] According to press reports, in 1996 Livan Hernandez signed for \$7.9 million, Rolando Arrojo for \$7 million, and Osvaldo Fernandez for \$3.2 million; in 2000 Adrian Hernandez signed for \$4 million; in 2001 Andy Morales signed for \$4.5 million; in 2003 Jose Contreras signed for \$32 million; in 2005 Kendry Morales signed for \$10 million (all figures in nominal dollars). These contracts generally involved some combination of a signing bonus and guaranteed salaries over a period of years that make them difficult to compare to typical draftees' contracts. See *Cuban Baseball Defectors* (2005).

[19] For example, Venezuela's Miguel Cabrera obtained a \$1.8 million signing bonus in 1999, at age 16; South Korea's Hee Seop Choi received a \$1.2 million bonus in 1999 at age 19; Japan's Ichiro Suzuki, a more-established professional, received a \$5 million signing bonus and a \$4 million salary in 2001, and the Seattle Mariners paid his former team \$13 million for the right to negotiate with him.

prices in this market so frequently escalate to and beyond the averages for first-round draftees hints at a possible over-bidding problem under competition.

There is also historical evidence of over-bidding by teams prior to the installation of the annual draft system in 1965. Indeed, the draft was baseball's third attempt to limit the size of bets teams were placing on risky young talent. From 1946-50 and again from 1953-57, the sport experimented with a "bonus baby rule."[20] At the conclusion of WWII, competition for players drove up signing bonuses to historically unprecedented levels, and team owners responded by requiring that players signed for a bonus above a certain amount (which varied from year to year but started at a relatively modest $4,000) must be carried on the signing team's major league roster for two years instead of being assigned to a minor-league apprenticeship. Since these "bonus babies" were generally not ready to contribute productively at the major league level, reserving space for them on the 25-man roster imposed a sizeable opportunity cost on signing teams; most were willing to carry no more than one or two at any one time. Nevertheless, there were reported bonuses of $75,000 in the late '40s (equivalent to over $500,000 in 1998 dollars) and $100,000-plus bonuses ($650,000 or more in 1998 dollars) were common in the '50s. It was also widely suspected that teams evaded the rule in various creative ways (including putting signees' relatives on the payroll). A new system was therefore adopted in 1957; it allowed bonus signees to be drafted by new teams if their signing teams did not protect them on an expanded 40-man roster at the end of their first year in professional baseball. Again, the hope was to limit the size and number of large bonuses teams doled out. But by 1964, the

[20] For an excellent review of the evolution of the baseball draft, see Simpson (2005), from whom this information is taken.

top bonus paid had reached $205,000 (roughly $1.08 million in 1998 dollars) and the current draft system was installed.

All else equal, of course, buyers will always prefer monopsony to competition. Attempts to suppress competitive bidding are not surprising. The relevant question here is whether competition had led to bids that were inconsistent with positive expected returns under most plausible assumptions about signees' legitimate prospects. There are hints that the answer is yes. In the early 1950s, when the "bonus baby" rule was allowed to lapse and bidding was truly competitive, there were many recorded bonuses of $100,000 or more. Since baseball revenues have risen far faster than the general level of prices, simply expressing bonuses from that period in real, inflation-adjusted terms probably does not convey the extent to which such bids stretched team's budgets. Given that the average big-league team grossed under $2 million in 1952, a $100,000 bonus represented more than five percent of revenue. By 1998, with average team revenue having grown to $82.6 million, a comparable (5.1 percent of gross) bonus would be $4.2 million, well above our estimated break-even bonus for first-round draftees, discussed earlier. In other words, unless there is reason to suppose that the recipients of the six-figure bonuses of the early '50s had significantly greater chances of success than the top picks of the '90s, we suspect teams often overbid for their services.

VI. Concluding remarks

In this study of the prices teams pay to sign ballplayers to their initial professional contracts, we have found intriguing evidence that decision-makers are prone to systematic errors when they must choose among a large number of alternatives and assign values to investments that may generate sizeable returns but have a low probability of doing so. This is not to say that the multi-million dollar bonuses teams pay their top draft picks are typically wasted. Despite the fact that over three-quarters of these "blue-chip" prospects are failures, those who succeed produce rents for their teams (in the form of marginal revenue products far in excess of their wages) that are sufficient to repay the bonuses paid to all draftees. As a benchmark, the dollars teams invest in the median first-round pick have an expected IRR of 33 percent.

Based on an examination of their behavior during the '90s, however, we conclude that teams allocate their investment funds inefficiently. Teams over-value high school players relative to college players: Over the sample period, the expected yield on high school players, who have longer apprenticeship periods than college players, was 27 percent, versus 43 percent for collegians. Teams over-value pitchers (24 percent expected annual return) relative to position players (41 percent). Teams fail to reduce bonuses sufficiently to the lower-quality players drafted in the second round (20 percent expected return) and the third round (13 percent). Smaller-market teams fail to aggressively exercise their monopsony power to negotiate lower bonuses commensurate with the lower revenues successful draftees will generate in their markets. Finally, when teams must bid competitively for talent (in cases where amateur players are exempt from the draft for one reason or another), bonus payments often soar to levels that are consistent with negative expected returns.

We submit that this evidence supports the view that research in cognitive psychology holds promise for enhanced understanding of the market mechanism. Critics of the experimental evidence on the winner's curse, prospect theory, and bounded rationality tend to be suspicious of studies of behavior under controlled conditions where stakes are low and there may be inadequate opportunities for learning by decision makers. Neither problem applies here. This market involves multi-million dollar decisions, and teams' interest in making them efficiently is reinforced by their desire to win on the field as well as maximize wealth. Further, baseball's annual draft is now 40 years old. Though we have found some evidence that teams learn from prior mistakes (e.g., by slowly reducing their propensity to draft high school players over the years), the magnitude and persistence of the inefficiencies we have identified suggest that they are not merely a by-product of ordinary "friction" in the process of equilibration, but symptomatic of deep-seated biases in decision making. Clearly, teams over-estimate their ability to identify successful prospects from the high school cohort, among pitchers, and as they reach deeper into the talent pool. Without greater access to teams' deliberative processes, we can only speculate about the precise reasons for such biases, but the behavioral literature provides an abundance of possibilities.

For example, teams' over-valuation of high school players may be a by-product of what Tversky and Kahneman (1986) have labeled *nontransparent dominance*. In an experimental setting, Tversky and Kahneman found that subjects asked to choose between two alternative "lotteries" had no difficulty identifying the one with the higher expected value when the uncertain payoffs were arranged so that subjects could narrow things down to the probabilities or payoffs that made the dominant choice transparent. But when the same choices were rearranged, making the comparison of probabilities and payoffs more complex, 58 percent of subjects chose the inferior alternative. Dominance can be, in the words of Tversky and Kahneman (1986, p. S265), "masked by a frame in which the inferior option yields a more favorable outcome in an identified state of the world." In their experiment, many subjects saw that two of three outcomes in the dominated alternative looked more favorable than in the dominant one—and failed to consider that the third outcome made all the difference. This is similar to the choice teams face between high school players and collegians, with the probabilities and payoffs shown in Table 4. It *appears* that high school draftees are more likely than collegians to become stars or good players, and that collegians merely have a better chance to become regulars. But, even apart from the statistical insignificance of all these differences (itself a possible cause of nontransparent dominance), we now know that this appearance is misleading—and that collegians have much higher expected returns because they arrive in the majors sooner.

We should note in closing, however, that there is considerable historical evidence that the participants in this market recognize the difficulties and potential pitfalls they face in choosing among so many highly uncertain alternatives. Indeed, the current draft system is actually major league baseball's third attempt to cope with teams' chronic tendency to over-value untried players. We have found that, in general, the draft does exactly that. The fact that the players' union has never attempted to significantly alter the system (despite its monopsonistic character) hints that the union shares the view that competitive bidding for amateur players might not enhance the sport's viability. If teams still make systematic errors in some ways in this market, they are at least rational enough to be looking for institutions to limit the damage from whatever root causes of misjudgment exist.

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Table 1

Summary statistics for first-round draft choices, 1990-97

Quality Category	Probability of Pick Achieving Quality Category (*p*)	Mean Number of Years to Reach Majors (t_0)	Mean Number of Marginal Wins per Season (*MP*)
Star	4.3%	2.22	8.05
Good	8.3%	2.88	4.63
Regular	14.0%	3.13	2.28

Table 2

Time profile of potential rents on successful draftees, by quality category

	t	NMRP	w	R
STAR	**t**	**NMRP**	**w**	**R**
(p=0.043)	2.22	7.41	0.3	7.106
	3.22	7.41	0.3	7.106
	4.22	7.41	1.777	5.629
	5.22	7.41	2.607	4.799
	6.22	7.41	3.614	3.792
	7.22	7.41	3.792	3.614
GOOD	**t**	**NMRP**	**w**	**R**
(p=0.083)	2.88	4.26	0.3	3.957
	3.88	4.26	0.3	3.957
	4.88	4.26	1.022	3.235
	5.88	4.26	1.498	2.758
	6.88	4.26	2.077	2.179
	7.88	4.26	2.179	2.077
REGULAR	**t**	**NMRP**	**w**	**R**
(p=0.14)	3.13	2.09	0.3	1.795
	4.13	2.09	0.3	1.795
	5.13	2.09	0.503	1.592
	6.13	2.09	0.737	1.357
	7.13	2.09	1.022	1.072
	8.13	2.09	1.072	1.022

Notes: p = probability of a draft pick achieving each quality class, t = time in years after initial signing, NMRP = net marginal revenue product, w = expected salary in millions, and R = expected rents in millions of dollars per year.

Table 3

Regression equations explaining variance in first-round bonus payments

	(1)	(2)	(3)	(4)
Year	0.119** (0.010)	0.120** (0.010)	0.120** (0.010)	0.121** (0.010)
Slot	-0.056** (0.005)	-0.056** (0.005)	-0.057** (0.005)	-0.056** (0.005)
Slot^2	0.0005** (0.0001)	0.0005** (0.0001)	0.0005** (0.0001)	0.0005** (0.0001)
College		-0.106** (0.037)	-0.105** (0.037)	-0.102** (0.037)
Pitcher			-0.004 (0.037)	
Market				-0.009 (0.008)
Adj. R^2	0.682	0.692	0.690	0.692

Notes: Dependent variable is the natural log of the real (1998 $) bonus. Standard errors in parentheses, ** indicates significance at the 1 percent level.

Table 4

Summary statistics for first-round high school and college draftees, 1990-97

High School Draftees

Quality Category	Probability of Pick Achieving Quality Category (*p*)	Mean Number of Years to Reach Majors (t_0)	Mean Number of Marginal Wins per Season (*MP*)
Star	4.8%	2.44	8.86
Good	10.3%	3.56**	5.16
Regular	12.4%	4.31**	2.07

College Draftees

Quality Category	Probability of Pick Achieving Quality Category (*p*)	Average Number of Years to Reach Majors (t_0)	Average Number of Marginal Wins per Season (*MP*)
Star	4.0%	1.97	7.11
Good	6.7%	1.84**	3.83
Regular	16.1%	2.25**	2.43

Notes: Z-tests on the college and high school probabilities (p) fail to reject equality of proportions for each of the three quality classes. For time deferment (t_0) and productivity (MP) t-tests of equality of means for the college/high school cohorts were conducted and ** indicates rejection at the 1 percent significance level.

Table 5

Summary statistics for pitchers and position players drafted in first round, 1990-97

Position	Probability of Pick Achieving Star Quality	Probability of Pick Achieving Good Quality	Probability of Pick Achieving Regular Quality
Pitchers	2.67%	2.67%**	16.67%
All Others	5.96%	13.91%**	11.26%

Note: Z-tests on the equality of pitcher and non-pitcher probabilities (p) were conducted for each quality class. ** indicates rejection at the 1 percent significance level for equality of proportions.

Table 6
Summary statistics for first three draft rounds

Draft Round	Median Real Bonus	Probability of Pick Achieving Star Quality	Probability of Pick Achieving Good Quality	Probability of Pick Achieving Regular Quality
First	$1.02M	4.3%	8.3%	14.0%
Second	$450k	0.9%*	2.2%**	6.3%**
Third	$270k	0.0%**	0.9%**	5.2%**

Notes: Median bonus for first round is based on data from 1995-2001; bonuses for second and third rounds are based on 1998 data. Probabilities are based on Callis (2003) survey covering 1990-'97 drafts. ** and * indicate rejection at the 1 and 5 percent significance levels respectively on a z-test for equality of proportions compared to first round.

Fig. 1: Rents Captured from a Successful Draftee:
The Case of "Chipper" Jones