Informed entry in auctions[1]

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Abstract

We examine entry decisions in first-price and English clock auctions with participation costs. Potential bidders observe their value and report maximum willingness to pay (WTP) to participate. Entry occurs if revealed WTP (weakly) exceeds the randomly drawn participation cost. We find no difference in WTP between auction formats, although males have a higher WTP for first-price auctions. WTP is decreasing in the number of potential bidders, but this reduction is less than predicted and small in magnitude.

JEL Classifications: C72, C90, D44, D80.

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1 Introduction

Auction environments where bidders face a costly entry decision abound. Individuals or firms have an opportunity cost of participating in an auction and may also face an explicit cost of preparing a bid (e.g., legal fees, financial modeling, consultants and advisers). These costs may influence the entry decisions of potential bidders. Yet, much of the auction literature studies the bidding behavior and revenue ranking of auctions with a fixed and exogenous set of bidders who do not face entry decisions or participation costs.

The main question this paper addresses is: Is entry into auctions sensitive to the auction format? This is an important question with practical implications for auction design. For instance, the revenue ranking of auction mechanisms depends, in part, on how many bidders each format is able to attract. If potential bidders are not indifferent among auction mechanisms, sellers should account for this. This is especially important when there is a small pool of potential bidders. Ex ante, the marginal expected value of a bidder to the auctioneer is highest when the expected number of bidders is low. One case where this occurs is when there are a small number of potential bidders.

If there are more potential bidders, do those with relatively low values opt not to participate in anticipation of a reduction in expected payoffs conditional on entry? Does any such change in entry behavior differ across auction formats? These questions are significant, since auction designers sometimes hold invitation-only auctions, thereby determining the number of potential bidders. Likewise, auction designers are often aware of the number of potential bidders. Does the optimal auction format depend on the number of potential bidders?

Theorists have informally speculated about the potential ways in which non-pecuniary preferences might influence bidder willingness to participate in various auction formats. For example, Engelbrecht-Wiggans (2001) suggests that oral or ascending auctions may be more attractive for bidders due to lower strategic uncertainty. Bidders in oral auctions may need or want to spend less effort acquiring and interpreting information than in sealed-bid auctions. Thus, it costs less to participate in oral auctions than in sealed-bid auctions. The lower participation cost could make oral auctions more attractive to bidders. On the other hand, Klemperer (2002) argues that ascending auctions are vulnerable to predatory behavior on the part of bidders, which might depress entry when there are even small participation costs.

This paper experimentally examines endogenous entry in independent private value auctions in which there is a cost of participation that is common to all potential bidders. Each bidder knows her value prior to her entry decision. We vary auction format between first-price and English clock on a between-subject basis, and vary the size of the pool of potential bidders on a within-subject basis.

We employ the Becker deGroot Marschak (BDM) procedure to elicit threshold entry decisions (i.e., maximum willingness to pay (WTP) to participate in the auction) (Becker et al., 1964). More precisely, in each auction, potential bidders are privately informed of their value, and then simultaneously report their maximum WTP to enter the auction, without knowing what the participation cost of the auction is. This common participation cost, which is randomly chosen, is then revealed to them. Those who reported a maximum WTP that (weakly) exceeds the chosen participation cost enter the auction and observe the number of bidders who entered before placing their bids. Given the complexity of using a BDM mechanism to determine participation in an auction, we relied on subjects who had previously participated in an experiment that involved direct entry decisions into auctions (Aycinena and Rentschler, 2014).[1]

Our focus is on environments where the seller is auctioning a unique good and/or does not directly compete with alternative mechanisms for potential bidders. Examples of such environments include real

[1] Aycinena and Rentschler (2014) uses within-subject variation on auction formats to evaluate revenue equivalence predictions between first-price and English clock auctions with endogenous entry and a fixed number of potential bidders, both when the bidders are informed regarding the number of entrants when choosing their bids, and when they are not. Our design differs from theirs in that we are able to more precisely measure entry behavior, and that we study the effect of varying the number of potential bidders.

estate auctions, art auctions, and a wide variety of government auctions (e.g., timber auctions, infrastructure procurement auctions, auctions for pollution permits and auctions to sell state-owned assets). We are not motivated by on-line auction environments where multiple sellers of homogeneous goods compete for bidders via their choice of auction format.

We find that WTP for both auction formats is increasing in the private value. We also find that reported WTP systematically exceeds equilibrium predictions and payoffs in both auction formats. This difference exists when there are three or five potential bidders and is consistent with other experiments on entry (Camerer and Lovallo, 1999; Fischbacher and Thoni, 2008). However, entry does not vary by auction format, although men are willing to pay more to enter a first-price auction. WTP is decreasing in the number of potential bidders, but this difference is small in magnitude and less than predicted by theory. We find that preferences for competition are partially able to explain why entry behaviors is less responsive to increased competition in the form of additional potential bidders.

The remainder of the paper is organized as follows. Section 2 summarizes the related literature. Section 3 provides theoretical predictions. Section 4 describes our experimental design. Section 5 contains results. Section **??** discusses the implications of our results and Section 6 concludes.

2 Related literature

Early theoretical analysis of endogenous entry in auctions has focused on cases where potential bidders observe the common cost of participation, and then decide whether or not to enter. Only after entry do potential bidders learn their type. Thus entry decisions are made before any private information is observed. Examples of this approach include McAfee and McMillan (1987), Levin and Smith (1994), Engelbrecht-Wiggans (1993) and Smith and Levin (1996).[2] Pevnitskaya (2004) generalizes this approach by allowing for heterogeneous levels of risk aversion. An individual's degree of risk aversion is private information which leads to a threshold entry decision based on risk preferences.[3]

More recently, the focus has been on environments in which potential bidders observed their value prior to entry. The branch of this theoretical literature that is most closely related to our design changes the timing of information revelation in the game. In particular, potential bidders observe the same commonly known entry cost and their independent private valuation of the good prior to deciding whether or not to enter the auction. As such, their entry decision is contingent on their valuation. Menezes and Monteiro (2000) was, to the best of our knowledge,the first to examine symmetric equilibrium for risk-neutral potential bidders in several auction formats in this environment. Cao and Tian (2010) analyzes this environment for first-price auctions. Lu (2009) examines optimal auction design when potential bidders observe their valuations prior to their entry decision, and all potential bidders have the same opportunity cost of entry.[4]

Despite the important theoretical progress, there has been relatively little empirical or experimental work on entry in auctions. The experimental literature can be divided in two main branches. The first branch of experimental literature focuses attention on the case in which bidders only learn their value after they have incurred their participation cost (Smith and Levin, 2002; Reiley, 2005; Ertaç et al., 2011).[5]

[2] In a related approach, Li and Zheng (2009) studies procurement auctions in which potential bidders only learn their private cost of supplying the good upon entering. This paper then tests the model using data from highway mowing auctions in Texas.

[3] A similar approach has examined entry with private information other than value before entry decision. Such private information includes participation costs independently drawn from a common distribution (Moreno and Wooders, 2011) or signals which provide information about the other valuations (Ye, 2004).

[4] Much attention in this environment has been on second-price auctions, as potential bidders who enter have a weakly dominant strategy to bid their valuation (e.g., Campbell 1998, Miralles 2008, Tan and Yilankaya 2006, Cao and Tian 2008). Green and Laffont (1984) and Cao and Tian (2009) allow both valuations and participation costs to be private information at the time of entry.

[5] Ertaç et al. (2011) measures willingness to pay to enter either a first or second price auction, and varies the number of opposing bidders and whether that is known before entry, after entry or not at all. Participating bidders face simulated opponents who bid

Engelbrecht-Wiggans and Katok (2005) report the result of experiments in which potential bidders must choose between participating in an auction or receiving an outside option. They elicit this choice for a range of possible outside options using a BDM mechanism. They found no statistically significant difference in the willingness to pay across auction formats, despite the fact that bidders earn significantly more in the ascending bid auction. Potential bidders over-enter into first-price auctions, but not ascending clock auctions. Their design differs from ours in several important ways. First, bidders in this experiment did not observe their value until after entry had occurred. That is, entry decisions were not able to be conditioned on bidder valuations. Second, bidders were not informed of the number of bidders when formulating their bids. Third, in their design, auction format was varied on a within-subject basis, whereas we vary this between subjects. Palfrey and Pevnitskaya (2008) reports the result of an experiment which demonstrates that bidders with relatively high degrees of risk aversion do self-select into a first-price auction. Further, there is an entry level effect, such that entry exceeds predictions.

The second branch of the literature focuses on the case in which multiple auction mechanisms selling identical goods compete for a fixed pool of bidders. Ivanova-Stenzel and Salmon (2011), for instance investigates an environment in which each bidder observes her value before making her entry decision. However, a fixed pool of bidders must choose which auction format to enter: a first-price or English clock auction. They find that revenue and efficiency are equal between the two formats. However, bidder payoffs are higher in the ascending bid auctions. To the best of our knowledge, this is the only other experimental examination of endogenous participation in independent private value auctions in which the bidder observes his value before making an entry decision. Our design differs from theirs in that, rather than being asked to choose between auction formats with a given value, bidders choose whether to participate or not in a given auction format by expressing their willingness to pay the participation cost.

This second branch of the experimental literature has received considerable attention (Ivanova-Stenzel and Salmon, 2004, 2008a,b; Engelbrecht-Wiggans and Katok, 2005). However, the focus differs from ours. We are interested in costly entry into auctions of goods without close substitutes where potential bidders decide whether or not to enter an auction, rather than environments in which multiple sellers complete via auction mechanism (as opposed to price). Their focus is on environments in which multiple auction mechanisms selling perfect substitutes compete for a fixed pool of bidders.[6] Notice that in these experiments, the focus is not on entry into auctions but on selection among auction formats. That is, subjects in such environments must enter an auction; the question is which format they will choose to enter. As such, entry behavior is not compared to an equilibrium prediction. Instead, relative payoffs between formats are compared to determine whether entry is too high relative to the other format.[7]

The paper that is closest to our is Aycinena and Rentschler (2014). In this paper, bidders make a binary decision regarding whether to enter a given auction format, after observing both their value and the common opportunity cost of entry. The number of potential bidders does not vary. Auction formats are varied within subject between first-price and English clock auctions. Whether or not bidders are informed of the number of participating bidders prior to formulating their bids is varied between subjects. They find that entry does not vary across auction format or information structure, although over-entry is observed in all environments. Their focus is on testing revenue equivalence predictions across their treatments. The binary entry choice that potential bidders face in this environment makes it difficult to measure entry thresholds with precision. That is the primary motivation for the design of the current paper. We wish to compare willingness to pay to enter both first-price and English clock auctions. We also wish examine how entry behavior depends on the number of potential bidders.

according to the risk-neutral equilibrium. They find significant over-entry in both auction formats, and across information structures.

[6]This may arise when multiple online auction sites sell identical products, but differ in the auction mechanism they employ.

[7]This implies that their results on entry are not comparable to our results. In their setting, over-entry in an auction format necessarily implies under-entry in another auction format. Our setting allows for over or under entry in either, both, or none of the auction formats.

3 Theory

A set of risk neutral players $\mathbf{N} \equiv \{1, ..., n\}$ are potential bidders in an auction for a single unit of an indivisible good. Each potential bidder $i \in \mathbf{N}$ privately observes her value of the good v_i, which is an independent draw of V, with distribution F and support $[0, v_H]$. The seller is presumed to have value of zero. There is a cost of participating in the auction, $c \in [0, c_H]$, which is common to all potential bidders. This cost, n and F are common knowledge. Upon entering the auction, all bidders are informed of the number of entrants, m, prior to choosing their bids.

In the unique symmetric equilibrium, potential bidders enter the auction only if their value weakly exceeds a threshold value, which we denote as v_c, for which they are indifferent regarding entry. Since equilibrium bid functions in the subsequent auction are monotonically increasing, a bidder with $v_i = v_c$ can only win the auction if she is the only entrant, which occurs with probability $F(v_c)^{n-1}$. In this case she obtains the good at a price of zero. Thus, the expected payoff of entering the auction with $v_i = v_c$ is $v_c F(v_c)^{n-1}$, and v_c must satisfy

$$v_c F(v_c)^{n-1} = c. \tag{1}$$

Notice that v_c does not vary by standard auction format. Since, in equilibrium, each bidder employs the same cutoff entry strategy, any bidder who has entered must have a valuation above v_c. Thus, the subsequent auction is a standard independent private value auction with m bidders in which each valuation is drawn from

$$F(v \mid v \geq v_c) = \frac{F(v_i) - F(v_c)}{1 - F(v_c)}. \tag{2}$$

In an English clock auction, bidders have a weakly dominant strategy to bid their value. As such, their equilibrium bid function is $\rho(v_i) = v_i$.[8] In a first-price auction (following Menezes and Monteiro (2000)) the equilibrium bid function is

$$\beta(v_i) = v_i - \left(\frac{1}{(F(v_i) - F(v_c))^{m-1}}\right) \int_{v_c}^{v_i} (F(t) - F(v_c))^{m-1} \, dt. \tag{3}$$

Menezes and Monteiro (2000) finds that first-price and English clock auctions in this environment are revenue equivalent. This expected revenue, R, is given by

$$R = n(n-1) \int_{v_c}^{v_H} (1 - F(t)) \, t F(t)^{n-2} f(t) \, dt. \tag{4}$$

That is, theory predicts that the revenue equivalence theorem generalizes to environments with endogenous entry. [9]

4 Experimental design

The primary objective of our experimental design is to examine entry thresholds between first-price and English clock auctions: Is entry behavior invariant across auction formats? We are further interested in testing whether observed entry corresponds to equilibrium predictions. Is the entry cost that potential bidders are willing to pay increasing in their value? Is the entry cost that potential bidders are willing to pay decreasing in the number of potential bidders?

[8] Derivations of equilibrium can be found in Appendix A.

[9] Aycinena and Rentschler (2014) tests revenue equivalence predictions between these two auction formats for a fixed set of potential bidders in an environment where potential bidders make binary entry decisions. Additionally, it evaluates whether or not informing bidders of the number of entrants prior to bidding affects revenue.

To investigate these questions, we employ a 2×2 experimental design that varies the number of potential bidders in a group within subjects and varies the auction format between subjects. In particular, in some sessions, the auction format is first-price, and in others it is English clock. Within a given session we alternate the number of potential bidders in a period between three and five in ten period blocks. That is, the number of potential bidders is constant for ten periods, and then is changed. A session has forty periods, so each subject participated in twenty periods with group size. In order to control for order effects, we vary the order in which subjects face these alternative group sizes. For each auction format, we ran a total of five sessions, three of which began with five potential bidders, and two of which began with three potential bidders. This design is summarized in Table 1.

An experimental session has fifteen participants and forty periods. In each period participants are randomly and anonymously matched into groups. Each group comprises a set of potential bidders for an auction. Values in each auction are independent draws from a uniform distribution on $\{0, ..., 100\}$. At the beginning of each period, potential bidder i observes her value (v_i) for that period, which is private information. The auction format for that period and the number of potential bidders are common knowledge. Likewise, each bidder knows that if she enters the auction she will be informed of the number of entrants (m) prior to choosing her bid.

In each period there is a common cost of entering the auction which is drawn from a uniform distribution on $\{1, ..., 30\}$, and is not initially observed by potential bidders. The cost of entry is restricted in order to reduce the number of auctions in which the cost of entry is so costly as to preclude any entry. In the first stage of a period, each potential bidder reports her WTP to enter the auction. Afterwards, all potential bidders are informed of the entry cost. If a potential bidder's WTP is at least as large as the entry cost, then she enters the auction, is informed of the number of entrants, and chooses her bid. Otherwise she does not enter the auction.[10] Reported WTP is restricted to be between zero and thirty-one. We opted to allow reported WTP to be either strictly smaller or strictly larger than the possible costs of entry so that participants would have an obvious way to indicate that they would like to enter or not regardless of the realized cost of entry.

This entry mechanism is an asset to our design because it allows us to obtain a much more precise measure of WTP than if potential bidders were simply asked to enter or not after observing their value and the entry cost.[11] However, participants may have a difficult time assessing these expected payoffs, since auctions are a complex environment. Indeed, Engelbrecht-Wiggans and Katok (2005) hypothesize that this drives their results in a similar experiment. Participants may also have a difficult time understanding that they maximize their utility by being truthful in the WTP elicitation.

Our design addresses both these concerns directly. To help ensure that a lack of understanding of auction environments with costly entry does not affect our results, we restricted to participants who had participated in a previous experiment (reported in Aycinena and Rentschler (2014)). In this previous experiment, there were forty-eight periods of auctions with endogenous entry, half of which were first-price and half of which were English clock. Potential bidders observed their values and the cost of entry, and then made a binary entry decision.

To ensure understanding of the BDM entry procedure in the current experiment, we carefully explained the procedure in the instructions, and provided examples which illustrated why being truthful was the optimal choice and why deviating from truthful revelation was weakly dominated. After the instructions, we tested understanding of the entry procedure prior to the experiment. In addition, after each entry decision in the experiment, a confirmation screen reminded participants of the implications of their decision contingent on the cost realization.[12] We address this concern such that we do not expect differences in behavior due

[10] That is, the Becker DeGroot Marschak (BDM) procedure (Becker et al., 1964) is used to elicit WTP, so that each participant has an incentive to report her true WTP.

[11] Alternatively, we could have elicited the minimum value they required to participate given the cost of entry. We choose to elicit WTP because such assessments are more likely to coincide with decisions faced by participants outside the lab.

[12] Specifically, the confirmation screen displayed what would happen if the participation cost were weakly less than their stated

to the entry procedure across treatments. Whether or not such differences exist in practice is an empirical question. We compare our data using the BDM mechanism for entry with that of Aycinena and Rentschler (2014) which uses binary entry. We find no evidence that BDM entry mechanism is biasing results.[13]

If a potential bidder does not enter the auction that period, she participates in a pastime while she waits for the auction to end. This pastime does not affect payoffs and is intended to mitigate boredom from inducing participants to report WTP in excess of their financial incentives. However, we also do not want the pastime to be so engaging as to reduce WTP. To this end, the pastime involves participants playing tic-tac-toe against a computer. Note that the pastime as well as the use of the BDM mechanism are consistent across both auction format and the number of potential bidders, so any treatment differences are not driven by the choice of pastime or elicitation mechanism on WTP.

Once the auction for that period has ended, each participant, regardless of whether or not they participated in the auction receives feedback. They are informed of the cost of entry, the number of bidders, the price at which the good was obtained (when applicable), all observed bids (ordered from highest to lowest), as well as their payoff for the period.[14]

All sessions were run at the Centro Vernon Smith Economía Experimental at Universidad Francisco Marroquín. Subjects were undergraduates of said institution. Table 2 provides summary statistics regarding the gender and age of subjects. The computer interface was programmed in z-Tree (Fischbacher, 2007). The software utilized the same realizations of values, entry costs and random re-matching of subjects in every session. Subjects were seated at computer terminals for the duration of the experiment. These terminals have dividers to prevent subjects from interacting outside of the computer interface. Once seated, subjects were shown video instructions (they were also provided with a hard copy of the instructions). This video contains screen shots of the computer interface in order to familiarize subjects with the environment. Once the video was completed, subjects were asked to complete a short quiz to ensure comprehension. Any remaining questions were then answered in private.

Each session lasted for approximately one and a half hours. Subjects were paid a $Q20 \approx US\$2.50$ show-up fee. All other monetary amounts in the experiment were denominated in experimental pesos ($E\$$), which were exchanged for Quetzales at a rate of $E\$7.5 = Q1$. Subjects began the experiment with a starting balance of five hundred experimental pesos to cover any losses. The average payoff was $Q84$, with a minimum of $Q36$ and a maximum of $Q147$. As a basis of comparison, the main employer of students on campus is the library which pays student workers $Q24$ per hour.

5 Results

The experiment is designed to allow us to examine entry thresholds across auction formats and number of potential bidders. Assuming equilibrium beliefs, a potential bidder's equilibrium WTP corresponds to the entry cost at which she is indifferent between participating or not. That is, WTP is predicted to satisfy $WTP = v_i \cdot F(v_i)^{n-1}$. For the parameters used, this is $WTP = v_i^n/100^{n-1}$, which ranges from zero to one hundred. However, in our experimental design entry costs are integers from one to thirty. In our analysis we define predicted WTP as the WTP a potential bidder is predicted to report given the constraints of our experimental design. That is, if the equilibrium WTP of a potential bidder is greater than thirty, predicted

WTP as well as what would happen if it was strictly greater, and asked them to confirm (or modify) their decision.

[13]For a detailed analysis comparing behavior between this two entry mechanisms, refer to Appendix C

[14]This is a relatively high level of feedback for auction environments. Neugebauer and Selten (2006) shows that providing losing bids as feedback may reduce overbidding in first-price auctions. Also, providing all observed bids as feedback in both auction formats could reduce the extent to which potential bidders prefer one format over the other. If potential bidders were to prefer English clock auctions due to an expectation of greater feedback, then providing all observed bids in first-price auctions may reduce their relative WTP for English clock auctions. Varying the level of feedback, as well as varying the relative level of feedback across auction formats, would be interesting questions for future research.

WTP is then thirty. Recall that although the maximum possible entry cost is thirty, we allow potential bidders to report WTP up to thirty-one so that there is a transparent way for them to always enter the auction regardless of the realized entry cost. In our analysis we censor all such observations to thirty.

5.1 Willingness to pay

Table 3 reports summary statistics for observed and predicted WTP, with both variables censored. Since the entry cost can never be smaller than one, potential bidders with a value such that $1 > v_i^n/100^{n-1}$ are predicted to report a WTP strictly less than one and never enter the auction. Likewise, since the entry cost cannot exceed thirty, potential bidders with a value such that $30 < v_i^n/100^{n-1}$ are predicted to report a WTP (weakly) greater than thirty, and always enter the auction.

We refer to the interval of values for which a potential bidder is never predicted to enter as region one. The interval of values for which predicted entry depends on the realized entry cost is referred to as region two, and the interval of values such that entry is always predicted is referred to as region three. When the number of potential bidders is three (five), region one consists of values strictly smaller than $22(40)$, region two is $22 \leq v_i \leq 66$ ($40 \leq v_i \leq 78$), and region three consists of values strictly above $66(78)$.

Before testing our main research question, we proceed to examine the data and how whether WTP corresponds to the equilibrium predictions. WTP is predicted to be an increasing function value, and that is precisely what we find: reported WTP is increasing in potential bidders' valuation. This is clearly illustrated in Figure 1, which shows mean reported and predicted WTP by valuation, group size and auction format.

Table 5 reports summary statistics for deviations of observed WTP from predicted WTP across auction format and group size, by region. Although mean reported WTP differs systematically from Nash predictions in regions one and three, this seems mainly due to the fact that in region one, subjects can only over-enter, and in region three they can only under-enter. This is a consequence of the restrictions our experimental design puts on reported WTP: subjects are not able to report WTP less than zero or greater than thirty-one. For region one, median reported WTP is one (three) when the number of potential bidders is three (five).[15] In both cases modal WTP is zero, as predicted by theory. For region three, modal and median reported WTP coincides with predicted WTP regardless of the number of potential bidders. The same cannot be said for region two, where the predicted decision is not trivial. Here, revealed WTP exceeds predictions regardless of auction format or group size (whether one looks at the mean or the median). In what follows, we focus our analysis on region two. Our results are generally robust to considering all regions. We refer the interested reader to the working paper version of this manuscript, which reports analysis using all regions pooled together as a robustness check.

Reporting WTP in excess of theoretical predictions results in over-entry in expectation. This is illustrated in Table 4, which contains summary statistics regarding the number of potential bidders that are predicted to enter the auction, as well as the associated predictions. Thus, we refer to this phenomenon as over-entry. Conversely, we refer to WTP below predictions as under-entry. As Table 3 and Table 5 show, we observe over-entry on average for all auction formats and group sizes in region two. A sign test using session level data to ensure independence of observations yields the same result in all four treatments: $z = 5$, $p = 0.031$. It is important to note that this observed over-entry is not an artifact of the BDM mechanism. Aycinena and Rentschler (2014) observe similar over-entry in environments in which potential bidders must make binary entry decisions. In fact, the observed level of over-entry using the BDM is smaller than that using binary entry decisions.

We further investigate the determinants of reported WTP. Table 6 presents random effects tobit estimates. We control for individual subject effects and account for the fact that reported WTP is censored. The dependent variable, WTP_{it}, is potential bidder i's reported WTP in period t. We include both a bidder's

[15] When there are five potential bidders, median WTP in region one drops to one for first-price auctions and two for English clock auctions in the second half of the experiment.

value at time t (v_{it}), as well the square of value as independent variables. We include treatment dummy variables (FP_i and G^5_{it}) and their interaction ($FP_i \cdot G^5_{it}$) to test for treatment effects. FP_i is our main variable of interest, as it allows us to test for auction format effects on entry. It takes the value of one when the auction format is first-price and zero otherwise. Similarly for G^5_{it}, it is one when there are five potential bidders, zero otherwise). Since the number of potential bidders was varied within-subject and we used different orders, we control for order effects. The dummy variable $GroupOrder_i$ takes the value of one if participant i began the experiment with a group size of five.[16] We also report specifications with additional controls: gender ($Male_i = 1$ if participant i is male), age (Age_i) and ($ln(t + 1)$) as a proxy for learning over the course of the experiment. Additionally, we report specifications that consider all forty periods, as well as specifications that restrict attention to the second half of the experiment.[17]

The effect of value on revealed WTP is positive and convex, as illustrated in Figure 1. To see this notice that the coefficients corresponding to v_{it} and v^2_{it} are both positive and significant. This is qualitatively in line with theory.

Moving to our central question, we find that as predicted by theory, entry behavior is invariant across auction formats, regardless of group size. The coefficients corresponding to FP_i and $FP_i \cdot G^5_{it}$ are small in magnitude, and never statistically significant. The first coefficient captures the effect of first price auctions with three potential bidders. The second captures the marginal effect of five potential bidders in first price auctions. The sum of the coefficients corresponding to FP_i and $FP_i \cdot G^5_{it}$ are never statistically different than zero at conventional levels. The p-values for tests of coefficients are reported below the relevant specification. These results are also in line with non-parametric tests.[18] It is important to note that subjects had previously participated in an experiment with endogenous entry in both of these auction formats, so inexperience regarding relative payoffs is unlikely to drive this result.

The invariance of entry across auction format is of interest because Ivanova-Stenzel and Salmon (2004) finds that, in an environment where bidders do not observe their value before making their entry decision, bidders have a higher WTP for English clock auctions.[19] Our results resemble those of Engelbrecht-Wiggans and Katok (2005), which finds no difference in WTP between first-price and English clock auctions. In our experiment however, bidders observe their value prior to entry and have previous experience in an experiment involving endogenous entry in auctions in which the auction format was varied (either first-price or English clock) on a within subject basis.

While we find revealed WTP is symmetric across auction formats, there is heterogeneity across gender: males report higher WTP for first-price auctions. In particular, although the coefficient on $Male_i$ is not significant, the interaction of $Male_i$ with FP_i is both positive and significant. This suggests that males prefer first-price auctions to English clock auctions, and that this preference is not shared by females.

We also find that revealed WTP is decreasing in the number of potential bidders. This is also qualitatively consistent with theory, but the magnitude is much smaller than predicted. On average, theory predicts that moving from three to five potential bidders should decrease WTP by 29%, but we observe a reduction of

[16] Recall that group size is either $n = 5$ or $n = 3$ for the first ten periods and then switches back and forth in ten period blocks.

[17] We also estimated models controlling for lagged entry, lagged win and lagged number of entrants. Lagged entry is not statistically significant. Lagged win is positive and statistically significant only during the second half. Lagged number of entrants is negative and statistically significant. The coefficients on the variables of interest are robust to including any or all of these controls. In the interest of brevity, we do not report these regressions here, although there are available on request.

[18] When there are three potential bidders there is no significant difference in WTP across auction formats when all regions are considered (robust rank order test, $\acute{U} = 1.448$, $p > 0.10$), as well as when attention is restricted to region two (robust rank order test, $\acute{U} = 0.669$, $p > 10.10$). Similarly, when there are five potential bidders there is no significant difference in WTP across auction formats in all regions (robust rank order test, $\acute{U} = 0.922$, $p > 0.10$), and when attention is restricted to region two (robust rank order test, $\acute{U} = 0.669$, $p > 0.10$). We rely on the p-values for the robust rank order test reported in Feltovich (2003).

[19] Ivanova-Stenzel and Salmon (2008b) and Ivanova-Stenzel and Salmon (2011) both find that bidders will often choose English clock auctions over first-price auctions when bidders do not know their valuation prior to entry. Since all else is equal between the two formats in their design, it is not possible to determine if this choice represents a higher WTP.

less than 5%.[20] The coefficient corresponding to G^5_{it}, which captures the effect of group size on WTP for English clock auctions, is negative in all specifications. However, when attention is restricted to the second half of the experiment, the standard errors increase and the coefficient is no longer significant. The effect on first-price auctions is captured by the sum of the coefficients corresponding to G^5_{it} and $FP_i \cdot G^5_{it}$. This sum is significant in all specifications, although only at the 10% level when all periods are considered.[21] When all regions are considered, the coefficient on group size is negative and statistically significant for all periods as well as for the second half. While we do find that the effect of group size is negative and statistically significant in our regressions, it is not always economically significant.

We next investigate how reported WTP compares with Nash predictions. Table 7 presents results of regressions similar to those in Table 6. Instead of controlling for value (linear and squared), we directly control for bidder i's predicted WTP during period t ($PWTP_{it}$). Note that $PWTP_{it}$ is a non-linear function of value and group size: $PWTP_{it} = v_{it} \cdot F(v_{it})^{n-1}$. Thus, the inclusion of this variable jointly tests whether or not WTP_{it} responds to changes in value and group size as predicted by theory. That is, it estimates the sensitivity of observed WTP to predictions. To determine whether there are level effects by treatment when controlling for $PWTP_{it}$ we include the dummy variables FP_i and G^5_{it}. We also interact these treatment dummies with $PWTP_{it}$. If entry behavior is consistent with theory on average, then the coefficient on $PWTP_{it}$ will equal one, and the remaining coefficients will not differ from zero. We also report specifications with additional controls: $Male_i$, Age_i, $ln(t+1)$, $GroupOrder_i$ and an interaction of gender and the first-price auction dummy. The first two specifications present results for all periods, both with and without the additional controls, and the last two specifications restrict attention to the second half of the experiment (last twenty periods).[22]

We find that reported WTP is increasing in predicted WTP, but this responsiveness is less than predicted by theory. To see this, note that in all specifications the coefficient corresponding to $PWTP_{it}$ is positive and significant, but is also significantly less than one.[23] In addition, the coefficient corresponding to the constant is positive and statistically significant, suggesting a level effect of over-entry. Also noteworthy is that there seems to be learning throughout the experiment, moving reported WTP closer to the theoretical prediction. To see this, note that, as in Table 6, the coefficient on $ln(t+1)$ is negative and statistically significant. Furthermore when we restrict attention to the second half of the experiment, the coefficient for $PWTP_{it}$ is higher and the constant is lower, suggesting greater responsiveness to theoretical predictions and a lower level of over-entry. The greater responsiveness to theoretical predictions in the second half holds even when controlling for $ln(t+1)$.[24]

Again, we find that both the intercept and the slope of reported WTP relative to theory do not vary by auction format. In particular, the coefficient for FP_i is not significantly different than zero in any specification. Furthermore, the sensitivity of WTP_{it} to predictions does not vary by auction format, since the coefficient corresponding to the interaction between $PWTP_{it}$ and FP_i is not statistically different from zero.

[20] This is on average across all regions, pooling the two auction formats. The average observed reduction in WTP is 5.66% for first price auctions and 4.01% for English clock auctions. Using non-parametric tests with session-level data across all regions and pooling across auction formats, we find that the difference in observed WTP across group size is marginally significant: sign test, $w = 8, p = 0.055$.

[21] The specifications reported separate the effect of group size by auction format. Although, in the interest of brevity, we do not report them here, we also consider specifications which pool the group size effect (by excluding the $FP_i \cdot G^5_{it}$ interaction). The pooled effect is significant in all specifications.

[22] As above, models controlling for lagged entry, win and number of entrants were also estimated. The results of these robustness checks mirror those corresponding to Table 6. Again, these regressions are not reported in the manuscript, although they are available on request.

[23] The p-values for tests of coefficients are reported below the relevant specification.

[24] This increase in responsiveness to theoretical predictions and the lower level of over-entry during the second half are statistically significant, although we do not report the relevant regressions in the interest of brevity.

The male preference for first-price auctions is also observed when controlling for $PWTP_{it}$. This result is illustrated in Figure 2, which compares WTP across treatments and gender in region two. Note that Figure 2 suggests that females may prefer English clock auctions, although this difference is not significant at conventional levels.

Interestingly, there is a large and significant level effect for group size. Sensitivity of WTP to theory is not affected by group size, as the coefficients corresponding to the interaction between $PWTP_{it}$ and G^5_{it} is not significant in any specification. However, the coefficient for G^5_{it} is positive and highly significant in all specifications. Further, the magnitude of these coefficients increases when we restrict attention to the second half. This does not mean that WTP increases with group size. What increases with group size is the deviation of observed WTP from predictions. While there is a slight reduction in observed WTP when the number of potential bidders increases, this reduction is much less than predicted by theory.

5.2 Payoffs

Given the two-stage nature of the game, it is possible that deviations from entry predictions in the first stage stem from expected non-equilibrium bidding in the second stage. Since (beliefs about) bidding behavior should only affect entry decisions through (beliefs about) payoffs, our results on WTP may be driven by expected payoffs in the subsequent auction. This is particularly true since potential bidders have experience in a similar experiment, and are thus likely to have an easier time forming accurate beliefs about relative payoffs than inexperienced participants.[25]

Table 8 presents summary statistics on bidder payoffs and WTP. Figure 3 illustrates the same exclusively for region two. We focus on payoffs in the auction, without the participation cost, as this is the relevant comparison to WTP. Payoffs are calculated based on the realized value and the observed number of entrants. We split the table into two panels. Panel A presents statistics with all three regions pooled. This allows us to make comparisons across different group sizes. However, this does not allow us to properly compare WTP to payoffs since WTP is truncated while payoffs are not. Thus, Panel B presents summary statistics only for region two. However, WTP in this region cannot be compared between group sizes. This is because for each group size, region two covers a different range of values.

The first two columns show observed and predicted WTP, conditional on entry. As both panels of the table illustrate, observed WTP exceeds predicted WTP for each treatment. Thus, we see that WTP exceeds ex-ante predicted payoffs for entrants. Since, the previous sub-section already analyzed WTP relative to theory, our focus here is on comparing WTP to payoffs.

The last two columns show observed and predicted auction payoffs conditional on entry, ignoring the (sunk) entry cost. These predicted auction payoffs take the observed entry as given, and calculate the expected payoff, assuming Nash bidding, of a bidder given what she knows after entry but prior to bidding: the observed number of bidders and her value.

Comparing the second and third columns in Panel B, we see that observed payoffs in the auction are lower than predicted WTP for all treatments. Thus, observed payoffs in the auction are lower than ex-ante predicted payoffs. However, when we condition predicted payoffs on the observed number of entrants, we find that bidders earn more than predicted. This can be seen by comparing the last two columns of the table. This difference is statistically significant and holds true for the pooled data (sign test, $w = 10$, $p = 0.001$), as well as by auction format, group size or auction format and group size (all these tests yield the same results: sign test, $w = 5$, $p = 0.031$).

[25] Recall that potential bidders have previous experience in a similar experiment, and are thus likely to have an easier time forming accurate beliefs about relative payoffs than inexperienced participants. This point is particularly important since Engelbrecht-Wiggans and Katok (2005) argues that experimental participants have a difficult time determining expected payoffs in a given auction format.

To understand why we see over-entry and observed payoffs that exceed predicted payoffs conditional on entry, one must look at bidding behavior upon entry. We find that on average auction participants underbid relative to predictions of Nash bidding conditional on number of entrants, holding beliefs about entry constant. On average, bids seem to fall somewhere in between Nash predictions conditional on the number of entrants, holding beliefs about entry constant, and a naive model in which bidders fall prey to the sunk cost fallacy. That is, where they subtract the (sunk) participation cost from their value. Appendix B presents a detailed analysis of bidding behavior.

So observed bidder payoffs exceed predicted payoffs conditional on entry. This, however, cannot justify the observed over-entry. This is because the magnitude of the difference between observed and predicted payoffs is not sufficiently large. We can see that in Panel B of Table 8. Note that the observed WTP of bidders considerably exceeds their observed payoffs in all treatments. This holds true for the pooled data (sign test, $w = 10$, $p = 0.001$), as well as analyzing by auction format, group size or auction format and group size together (all of these tests yield the same results: sign test, $w = 5$, $p = 0.031$). Not only are participants over-entering with respect to theory, but with respect to realized payoffs as well. Such over-entry relative to payoffs is not unique to our data. It has been documented in other experiments on entry such as Palfrey and Pevnitskaya (2008); Fischbacher and Thoni (2008).

Although observed payoffs in the auction cannot explain the level of over-entry, a more relevant question is to examine how observed payoffs are related to WTP across treatments. Theory assumes that a potential bidders WTP is determined by expected payoffs in the auction. Expected auction payoffs across treatments are correlated with observed payoffs. Thus, we would expect to find that payoffs are lower when group size is larger, and that payoffs do not vary across auction formats. As panel A shows of Table 8 shows, payoffs are greater for groups of three than for groups of five: on average, the difference in payoffs is 10.69 for first price auctions and 12.70 for English clock auctions. These differences are statistically significant at the session level when the data is pooled: sign test, $w = 10$, $p = 0.001$. The same result holds when considering each auction format separately: sign test, $w = 5$, $p = 0.031$.

Across auction formats, we find that payoffs are greater in English clock auctions. Using the robust rank order test for the data pooled across group sizes, we get $\acute{U} = -1.768$ with $p < 0.038$. This difference, however, is not very robust, as it is not statistically significant when restricting to either group size. Furthermore, the difference in payoffs across auction formats is small in magnitude. For five potential bidders the difference is 2.05, while for three potential bidders it is 4.07.

Turning our attention to gender differences, we found above that males reported a higher WTP for first-price auctions. One possible explanation is that their payoffs are higher in this auction format. However, payoffs for males are actually lower in first-price auctions than in English clock auctions, although this difference is not significant (sign test, $w = 9$, $p = 0.623$). As such, relative payoffs are also unable to explain the male preference for first-price auctions. Further, while males do earn more than females in first-price auctions this difference is not statistically significant (sign test, $w = 5$, $p = 0.187$). This result also holds for the pooled sample (sign test, $w = 7$, $p = 0.172$).

6 Conclusion

We experimentally examine threshold entry decisions in independent private value auctions where participation is costly and bidders learn their value before they make their entry decisions. In particular, we elicit WTP using the BDM mechanism. Once each bidder has reported her WTP, a participation cost that is common to all potential bidders is drawn from a discrete uniform distribution. If reported WTP (weakly) exceeds this participation cost then the bidder incurs this cost and enters the auction. Bidders are then told how many bidders there are in the auction and then place their bids.

We vary the auction format between a first-price auction and an English clock auction on a between-subject basis. In addition, we vary the size of the pool of potential bidders between three and five, on a within-subject basis.

We find that, consistent with theoretical predictions, entry threshold strategies are invariant to auction format. That is, reported WTP does not, on average, vary across auction formats; neither in level nor in sensitivity to theoretical predictions. This is relevant information for practical auction design. However, there seems to be heterogeneity in preferences across gender, since reported WTP for males is higher for first-price auctions. This finding cannot be explained by greater male profits in first price auctions. We should take this heterogeneity across preferences with a grain of salt, since this was not a hypothesis we originally planed to test and the result is only marginally significant during the second half of the experiment.

We also find that WTP is much less responsive than predicted to changes in the number of potential bidders. Despite the fact that both expected and observed payoffs are decreasing in group size, potential participants reduction in WTP to increases in group size by less than $1/6$ of the predicted reduction. This also has practical implications for auction design. Efforts to increase the group size seem to have little downside, something to auctioneers may want to keep in mind for invitation only auctions.[26]

In accordance with theory, we find that WTP is increasing in bidder valuation. However, entry thresholds under-respond to theoretical predictions. For the region of values where theory predicts that the entry decision will depend on the realization of the entry cost, we find that WTP increases at about $2/3$ of what theory predicts ($3/4$ when looking at the second half). In addition, there is a level increase in reported WTP regardless of value. This results in, on average, higher than predicted WTP.

When bidders report WTP in excess of predicted WTP, we say that they are over-entering the auction, because, on average, they will enter the auction more than predicted by theory. The observed over-entry exists in both auction formats, and persists throughout the experiment. This is despite that fact that they are paying more to enter the auction than they earn, on average. It should be noted that behavior is moving in the directions of equilibrium predictions. Sensitivity of observed WTP to predicted WTP increases and the level effect of over entry decreases during the second half.

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7 Tables

Table 1: Summary of experimental design

Potential bidders sequence	First-price	English clock
3535	2	2
5353	3	3

Notes: Table contains the number of sessions in each environment. There were fifteen subjects in each session. If a session started with ten periods with three (five) potential bidders, followed by ten periods with five (three) potential bidders, and so on, the potential bidder sequence is denoted as 3535 (5353).

Table 2: Summary statistics of the subjects, by auction format in session

Auction format	Percentage female	Age
First-price	0.36	21.11
	(0.48)	(2.12)
English clock	0.43	20.72
	(0.50)	(2.17)

Notes: Table contains means with standard deviations in parentheses.

Table 3: Summary statistics for observed (by auction format) and predicted WTP, for each group size

Groups size	Observed FP	Observed EC	Predicted
3 potential bidders	15.95	15.26	14.89
	(11.72)	(11.62)	(12.71)
5 potential bidders	15.04	14.65	10.56
	(11.75)	(11.44)	(12.47)

Notes: Table contains means with standard deviations in parentheses. Predicted willingness to pay is invariant across auction formats and it is based on the actual values drawn from the uniform distribution.

Table 4: Summary statistics for observed (by auction format) and predicted number of bidders, for each group size

Groups size	Observed FP	Observed EC	Predicted
3 potential bidders	1.69	1.60	1.54
	(0.92)	(0.92)	(0.86)
5 potential bidders	2.62	2.53	1.71
	(1.29)	(1.31)	(1.02)

Notes: Table contains means with standard deviations in parentheses. The predicted number of bidders is invariant across auction formats and it is based on the actual values drawn from the uniform distribution.

Table 5: Summary statistics for WTP deviations from theoretical prediction, by region and treatment

	Groups size	First-price	English c lock
Region 1			
	3 potential bidders	3.89	3.33
		(7.40)	(6.12)
	5 potential bidders	5.76	5.51
		(8.09)	(7.23)
Region 2			
	3 potential bidders	3.54	2.76
		(9.81)	(9.22)
	5 potential bidders	7.92	7.42
		(9.79)	(9.92)
Region 3			
	3 potential bidders	-3.80	-4.42
		(6.50)	(7.03)
	5 potential bidders	-3.39	-3.84
		(6.17)	(6.40)

Notes: Table contains means with standard deviations in parentheses of the difference between the observed and the predicted willingness to pay (based on the actual values drawn from the uniform distribution).

Table 6: WTP estimates for region two

	All 40 periods		Last 20 periods	
	(1)	(2)	(3)	(4)
v_{it}	0.196**	0.193**	0.233**	0.231**
	(0.077)	(0.077)	(0.112)	(0.112)
v_{it}^2	0.002***	0.002***	0.003**	0.003**
	(0.001)	(0.001)	(0.001)	(0.001)
FP_i	1.016	-3.143	0.448	-2.893
	(1.467)	(2.232)	(1.714)	(2.612)
G_{it}^5	-0.979*	-1.048*	-1.009	-1.305
	(0.552)	(0.551)	(0.806)	(0.812)
$FP_i \cdot G_{it}^5$	0.032	0.056	-0.806	-0.747
	(0.713)	(0.711)	(1.034)	(1.030)
$GroupOrder_i$	-0.374	-0.572	-0.904	-0.965
	(1.463)	(1.416)	(1.686)	(1.656)
$ln(t+1)$		-0.534**		-3.653**
		(0.241)		(1.462)
$Male_i$		-2.736		-1.959
		(1.970)		(2.300)
$FP_i \cdot Male_i$		7.227**		5.892*
		(2.826)		(3.297)
Age_i		-0.751**		-0.776**
		(0.327)		(0.380)
$Constant$	-0.775	18.180**	-2.545	27.396***
	(2.274)	(7.284)	(3.094)	(9.963)
Tests of coefficients[a]				
$FP_i + FP_i \cdot G_{it}^5 = 0$	0.481	0.168	0.838	0.166
$G_{it}^5 + FP_i \cdot G_{it}^5 = 0$	0.088	0.074	0.025	0.012
$Male_i \cdot FP_i + Male_i = 0$		0.027		0.096
$Male_i \cdot FP_i + FP_i = 0$		0.023		0.157
Observations	2,494	2,494	1,292	1,292
Left-censored observations	226	226	147	147
Right-censored observations	441	441	252	252
Log-likelihood	-7,188.8	-7,179.9	-3,646.0	-3,639.0
Bayesian information criterion	14,447.9	14,461.5	7,356.6	7,371.0
Akaike's information criterion	14,395.5	14,385.8	7,310.1	7,303.9

Notes: Determinants of entry cutoff strategies for region 2, where predicted WTP is uncensored. Estimates using individual level random effects tobit specifications. Dependent variable is reported willingness to pay to enter the auction. Standard errors reported in parentheses.

$^{+}p < 0.10, ^{*}p < 0.05, ^{**}p < 0.01, ^{***}p < 0.001$

[a] p-values for each relevant specification are reported in the corresponding column.

Table 7: Responsiveness of WTP to theoretical predictions for region two

	All 40 periods		Last 20 periods	
	(1)	(2)	(3)	(4)
$PWTP_{it}$	0.653***	0.648***	0.748***	0.754***
	(0.039)	(0.039)	(0.058)	(0.058)
$PWTP_{it} \cdot FP_i$	0	0	0.058	0.060
	(0.045)	(0.045)	(0.066)	(0.066)
FP_i	1.028	-3.161	-0.449	-3.847
	(1.499)	(2.251)	(1.771)	(2.653)
$PWTP_{it} \cdot G^5_{it}$	0.022	0.024	-0.055	-0.063
	(0.046)	(0.045)	(0.066)	(0.066)
G^5_{it}	5.920***	5.796***	6.888***	6.723***
	(0.560)	(0.561)	(0.813)	(0.812)
$ln(t+1)$		-0.596*		-3.681*
		(0.244)		(1.469)
$GroupOrder_i$		-0.613		-1.223
		(1.417)		(1.660)
$Male_i$		-2.712		-1.866
		(1.972)		(2.305)
$FP_i \cdot Male_i$		7.284*		5.983+
		(2.830)		(3.305)
Age_i		-0.758*		-0.792*
		(0.328)		(0.381)
$Constant$	5.742***	25.168***	4.737***	35.667***
	(1.094)	(7.067)	(1.314)	(9.642)
Tests of coefficients[a]				
$PWTP_{it} = 1$	0.000	0.000	0.000	0.000
$PWTP_{it} + PWTP_{it} \cdot FP_i = 1$	0.000	0.000	0.001	0.001
$FP_i + PWTP_{it} \cdot FP_i = 0$	0.489	0.159	0.823	0.151
$PWTP_{it} + PWTP_{it} \cdot G^5_{it} = 1$	0.000	0.000	0.000	0.000
$Constant = 0$	0.000	0.000	0.000	0.000
$Male_i \cdot FP_i + Male_i = 0$		0.024		0.083
$Male_i \cdot FP_i + FP_i = 0$		0.024		0.323
Observations	2,494	2,494	1,292	1,292
Left-censored observations	226	226	147	147
Right-censored observations	441	441	252	252
Log-likelihood	-7208.2	-7198.7	-3650.3	-3642.8
Bayesian information criterion	14479.0	14499.0	7357.9	7378.7
Akaike's information criterion	14432.4	14423.4	7316.6	7311.6

Notes: Estimated entry cutoff strategies relative to theoretical predictions for region 2, using individual level random effects tobit specification. Dependent variable is reported willingness to pay to enter the auction. Standard errors reported in parentheses.

$^{+}p < 0.10, ^{*}p < 0.05, ^{**}p < 0.01, ^{***}p < 0.001$

[a] p-values for each relevant specification are reported in the corresponding column.

Table 8: Summary statistics for WTP and payoffs, taking observed entry as given

Treatment	Observed WTP	Predicted WTP	Observed auction payoff	Predicted auction payoff
Panel A: All regions				
FP 3	24.056	21.153	21.572	15.214
	(7.930)	(11.235)	(29.348)	(22.914)
FP 5	23.861	16.826	10.880	4.744
	(8.003)	(12.592)	(20.604)	(13.901)
EC 3	23.903	21.543	25.638	11.588
	(7.648)	(11.114)	(31.720)	(24.639)
EC 5	23.637	17.206	12.934	4.579
	(7.581)	(12.627)	(24.783)	(16.286)
Panel B: Region Two				
FP3	19.953	11.868	9.837	5.934
	(8.601)	(8.140)	(18.544)	(13.499)
FP5	22.907	11.441	5.86	2.224
	(7.472)	(8.688)	(14.944)	(9.828)
EC3	19.823	12.173	10.58	5.732
	(8.109)	(8.281)	(19.929)	(14.395)
EC5	22.925	11.532	6.937	2.928
	(6.901)	(8.833)	(18.587)	(11.898)

Notes: Table contains means with standard deviations in parentheses. The first two columns contain observed and predicted willingness to pay, conditional on entry. The next two columns show observed and predicted payoffs for entrants, ignoring the sunk participation cost. Predictions are calculated based on the realized value. Predicted auction payoffs are the expected payoff of a bidder, assuming Nash bidding, conditional on what she knows after entry but prior to bidding: her value and the observed number of entrants. Panel A presents pooled data across all regions. Panel B restricts to data in region two. It should be noted that region two covers a different range of values for groups of three versus groups of five.

Table 9: Random effects tobit estimates of WTP relative to theoretical predictions for region two, controlling for competitiveness

	(1)	(2)	(3)	(4)
$PWTP_{it}$	0.666***	0.666***	0.670***	0.666***
	(0.043)	(0.043)	(0.043)	(0.043)
$PWTP_{it} \cdot FP_i$	-0.007	-0.007	-0.014	-0.007
	(0.050)	(0.050)	(0.050)	(0.050)
FP_i	-2.655	-3.6	-3.584	-1.064
	(2.623)	(2.657)	(2.665)	(3.723)
$PWTP_{it} \cdot G^5_{it}$	0.031	0.031	0.029	0.031
	(0.050)	(0.050)	(0.050)	(0.050)
G^5_{it}	5.956***	5.957***	3.909***	5.951***
	(0.620)	(0.620)	(0.895)	(0.620)
$ln(t+1)$	-0.193	-0.196	-0.17	-0.196
	(0.269)	(0.269)	(0.269)	(0.269)
$GroupOrder_i$	-0.26	-0.376	-0.347	-0.405
	(1.642)	(1.624)	(1.630)	(1.618)
$Male_i$	-1.864	-2.426	-2.443	-2.867
	(2.350)	(2.348)	(2.356)	(2.382)
$FP_i \cdot Male_i$	6.527*	7.567*	7.623*	7.746*
	(3.337)	(3.346)	(3.357)	(3.337)
Age_i	-0.925*	-0.929*	-0.933*	-0.863*
	(0.430)	(0.425)	(0.426)	(0.428)
$Comp_i$		0.045	0.026	
		(0.028)	(0.028)	
$Comp_i \cdot G^5_{it}$			0.044**	
			(0.014)	
$FP_i \cdot Comp_i$				0.02
				(0.038)
$EC_i \cdot Comp_i$				0.074+
				-0.041
$Constant$	26.436**	24.858**	25.742**	22.461*
	-8.975	-8.915	-8.948	-9.214
Tests of coefficients[a]				
$Comp_i + Comp_i \cdot G^5_{it} = 0$			0.016	
$Comp_i \cdot FP_i = Comp_i \cdot EC_i$				0.333
Observations	2,076	2,076	2,076	2,076
Left-censored observations	195	195	195	195
Right-censored observations	387	387	387	387
Log-likelihood	-5,912.3	-5,911.0	-5,906.0	-5,910.6
Bayesian information criterion	11,924.0	11,929.0	11,926.7	11,935.7
Akaike's information criterion	11,850.7	11,850.1	11,842.1	11,851.1

Notes: Estimated entry cutoff strategies (WTP) relative to theoretical predictions for region 2, using individual level random effects tobit specification. Estimation is restricted to sub-sample with available data on competitiveness measure from separate experimental task. Dependent variable is reported willingness to pay to enter the auction. Standard errors reported in parentheses.

$^+p < 0.10, ^*p < 0.05, ^{**}p < 0.01, ^{***}p < 0.001$

[a] p-values for each relevant specification are reported in the corresponding column.

Table 10: Random effects tobit estimates of WTP relative to theoretical predictions for region two controlling for risk preferences

	(1)	(2)	(3)
$PWTP_{it}$	0.657***	0.656***	0.656***
	(0.041)	(0.041)	(0.041)
$PWTP_{it} \cdot FP_i$	-0.015	-0.012	-0.012
	(0.048)	(0.048)	(0.048)
FP_i	-3.396	-3.833+	-3.375
	(2.366)	(2.308)	(5.267)
$PWTP_{it} \cdot G_{it}^5$	0.02	0.019	0.019
	(0.048)	(0.048)	(0.048)
G_{it}^5	5.793***	5.794***	5.794***
	(0.590)	(0.590)	(0.590)
$ln(t+1)$	-0.711**	-0.708**	-0.708**
	(0.258)	(0.258)	(0.258)
$GroupOrder_i$	-1.309	-0.632	-0.61
	(1.463)	(1.442)	(1.460)
$Male_i$	-2.523	-3.157	-3.12
	(2.031)	(1.988)	(2.023)
$FP_i \cdot Male_i$	6.386*	6.869*	6.828*
	(2.944)	(2.868)	(2.899)
Age_i	-0.752*	-0.755*	-0.758*
	(0.327)	(0.319)	(0.320)
$SafeChoices_i$		-1.436**	
		(0.497)	
$FP_i \cdot SafeChoices_i$			-1.474*
			(0.635)
$EC_i \cdot SafeChoices_i$			-1.374+
			(0.812)
$Constant$	25.956***	32.249***	31.997***
	(7.016)	(7.192)	(7.649)
Tests of coefficients[a]			
$SafeChoices_i \cdot FP_i = SafeChoices_i \cdot EC_i$			0.923
Observations	2,237	2,237	2,237
Left-censored observations	205	205	205
Right-censored observations	385	385	385
Log-likelihood	-6,487.0	-6,482.9	-6,482.9
Bayesian information criterion	13,074.2	13,073.8	13,081.5
Akaike's information criterion	13,000.0	12,993.8	12,995.8

Notes: Estimated entry cutoff strategies (WTP) relative to theoretical predictions for region 2, using individual level random effects tobit specification. Estimation is restricted to sub-sample with available data on risk preferences from separate experimental task. Dependent variable is reported willingness to pay to enter the auction. Standard errors reported in parentheses.
$^+p < 0.10, ^*p < 0.05, ^{**}p < 0.01, ^{***}p < 0.001$
[a] p-values for each relevant specification are reported in the corresponding column.

8 Figures

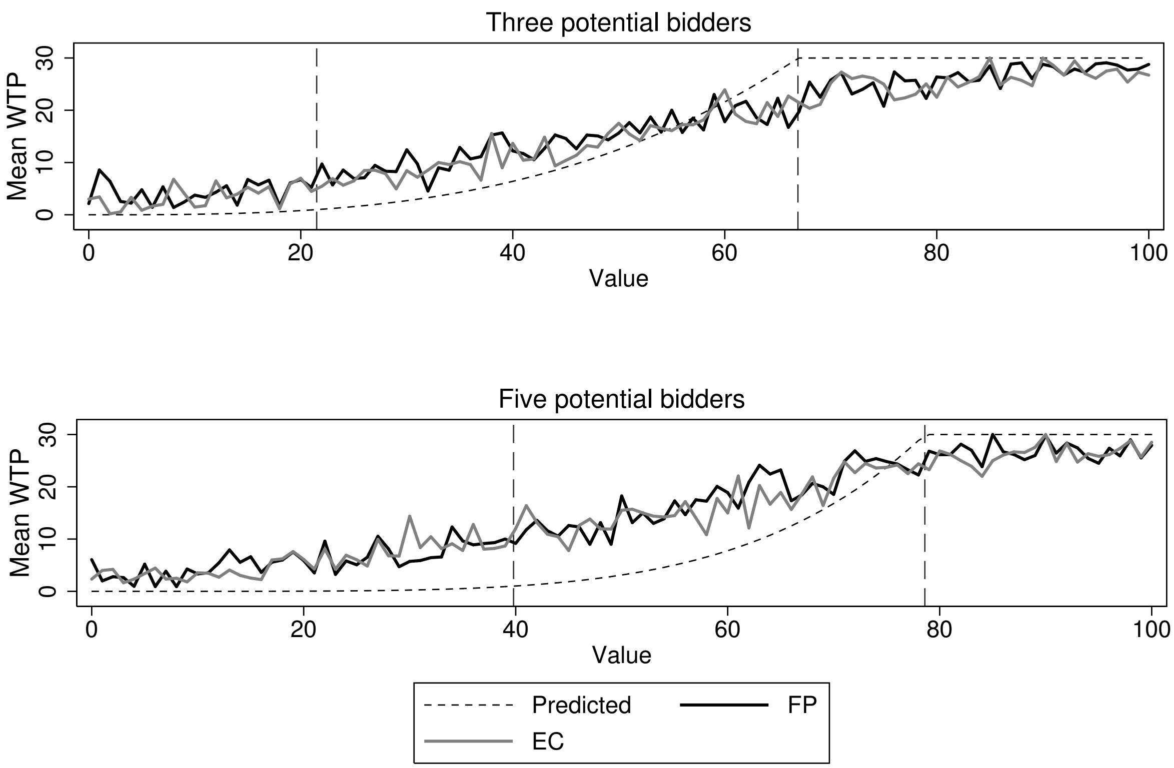

Figure 1: Mean WTP by value and number of potential bidders.

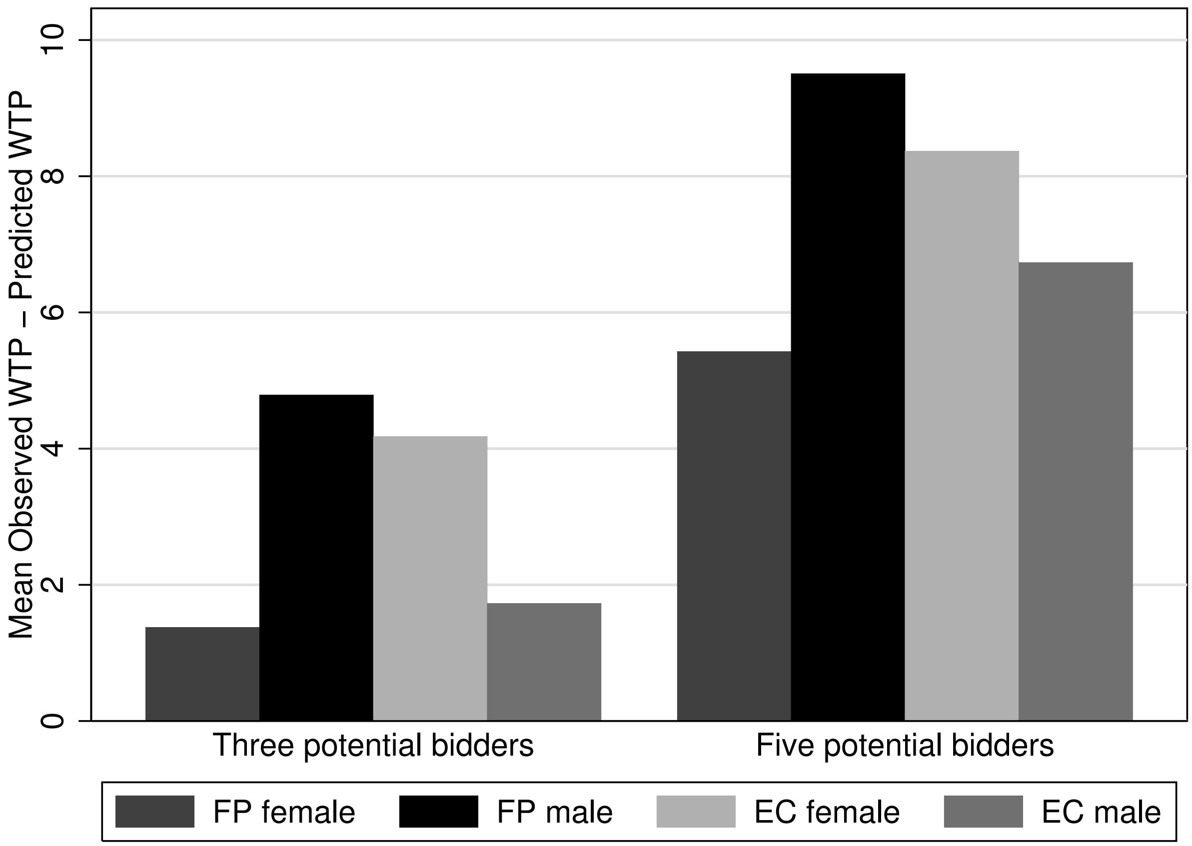

Figure 2: Deviations of WTP from Nash predictions in region two by gender.

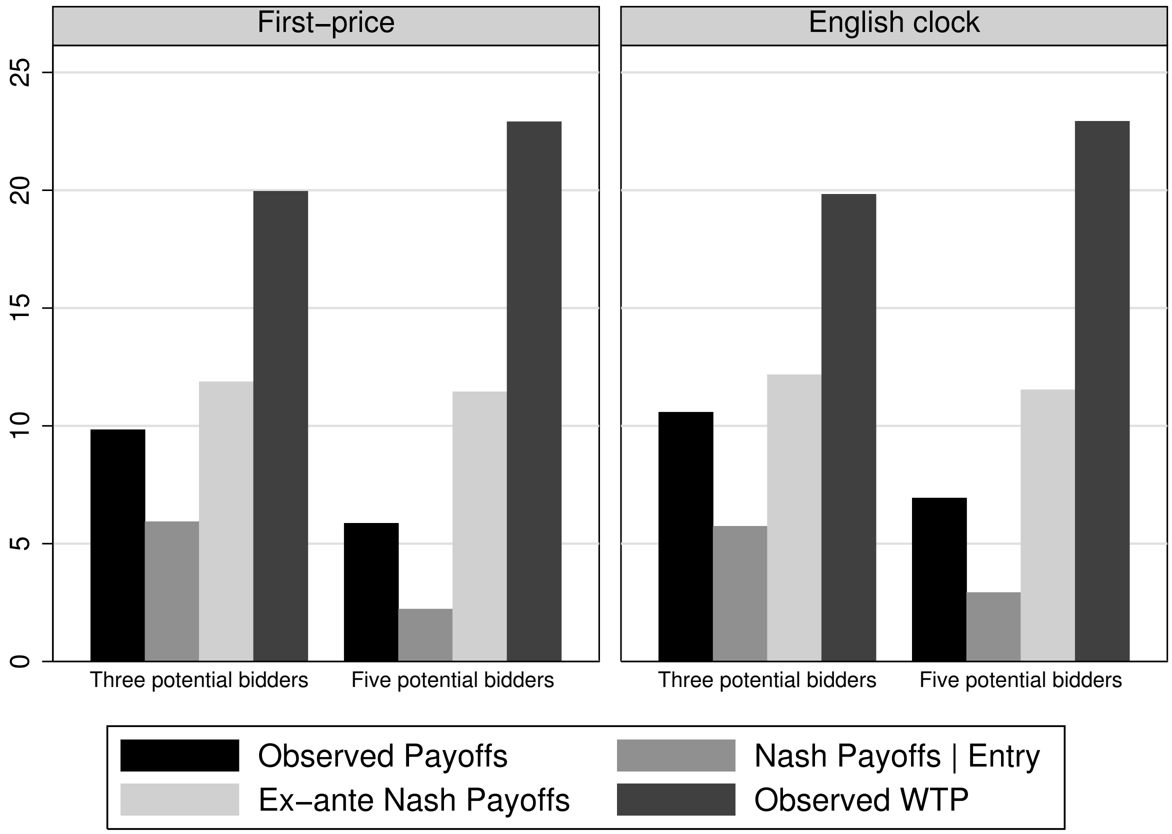

Figure 3: Payoffs by auction format and number of potential bidders in region two.

A Derivations of Nash predictions

A.1 First-price auctions

There are n potential bidders of which $2 \leq m \leq n$ have entered the auction. Assume that m is common knowledge. We restrict attention to symmetric equilibrium in which potential bidders employ an entry threshold of $0 < v_\omega < v_H$. Thus, all m bidders have values in excess of this threshold. This is then a standard first-price auction with independent private values. We refer the reader to Menezes and Monteiro (2000) for detail derivations. The equilibrium bid function for the m bidders is then given by

$$\beta(v_i) = v_i - \frac{\int_{v_\omega}^{v_i} (F(t) - F(v_\omega))^{m-1} dt}{(F(v_i) - F(v_\omega))^{m-1}}. \tag{5}$$

Now consider the case in which $m = 1$. That is there is only one bidder, and this bidder is aware of this. She will thus submit a bid of zero, and obtain the good.

Plugging the equilibrium bid function into the payoff function shows that the equilibrium payoff of a bidder with value $v_i \geq v_\omega$ and $m > 1$ is given by

$$\pi_i^{FP}(\beta(v_i), v_i|m) = \int_{v_\omega}^{v_i} \left(\frac{F(t) - F(v_\omega)}{1 - F(v_\omega)}\right)^{m-1} dt. \tag{6}$$

Note that if $m = 1$, then the equilibrium payoff of bidder i is simply v_i.

If a potential bidder has $v_i = v_\omega$, then she will enter the auction. She will then obtain the good at a price of zero, provided she is the only entrant, which occurs with probability $F(v_\omega)^{n-1}$. Setting the associated expected payoff equal to c implicitly defines v_ω:

$$v_\omega F(v_\omega)^{n-1} = c. \tag{7}$$

The ex ante expected revenue generated by the auction is

$$R_{FP} = n(n-1)\int_{v_\omega}^{v_H} (1 - F(t))\, tF(t)^{n-2} f(t)\, dt. \tag{8}$$

A.2 English Clock auctions

Assume that potential bidders employ a symmetric entry threshold, which we denote as $0 < v_\theta < v_H$. Note that any bidder who enters will always bid her value, regardless of the number of bidders, m. That is, the bidding function that is consistent with a perfect Bayesian equilibrium is $\rho(v_i) = v_i$. Thus, we only need to determine value of v_θ. Logic identical to the first-price auction shows that v_θ is implicitly defined by

$$v_\theta F(v_\theta)^{n-1} = c. \tag{9}$$

Note that the equilibrium entry thresholds are identical in first-price and English clock auctions. Since this implies that the expected payoffs of potential bidders are also identical, the expected revenue between the two formats is also identical. That is, $R_{EC} = R_{FP}$.

B Bidding behavior

Potential bidders enter more often than predicted. If bidders were reporting a WTP in keeping with their beliefs regarding their expected earnings, the observed over-entry would imply that they believe that bids are, on average, below the Nash prediction, thus increasing expected payoffs of entering the auction.[27] Since in English clock auctions bidders have a weakly dominant strategy to bid their value, potential bidding holding such beliefs seems unlikely. In first-price auctions equilibrium bids depend on both the observed number of bidders, as well as beliefs about the entry threshold employed by bidders. In particular, the Nash bidding function in first-price auctions is linear in the bidder's value. The slope of this function, $(m-1)/m$, is increasing in the number of bidders. The intercept, v_c/m is decreasing in the number of bidders and increasing in the cutoff entry value.[28] The entry cost has already been incurred at the bidding stage. In an English clock auction, it is a sunk cost that does not affect the weakly dominant bidding strategy. In a first price auction, it is still a sunk cost, but in equilibrium, it provides information on the minimum value of entrants. As such, it affects the minimum bid independent of value.

Bidding behavior relative to Nash predictions is illustrated in Figure 4. Summary statistics are contained in Table 11. Notice that bidding in first-price auctions is bimodal, one of which represents overbidding, the other of which represents underbidding. For English clock auctions we see a substantial amount of bidding in accordance with theory, as well as some underbidding. Significantly, note that in all treatments except FP5, bidding is, on average, below Nash predictions. This result is puzzling, since there is a large literature which finds that bidders in English clock auctions learn quickly to bid their value (Harstad, 1990) and that bidders in first-price auctions tend to overbid (Kagel and Levin, 1993). Many possible explanations for overbidding in first-price auctions. Explanations related to the preferences of bidders include risk averse bidders (Cox et al., 1983, 1988), a joy of winning Cox et al. (1992); Holt and Sherman (1994), and regret averse bidders (Engelbrecht-Wiggans and Katok, 2007; Filiz-Ozbay and Ozbay, 2007; Engelbrecht-Wiggans and Katok, 2009). Explanations which relax the assumptions of Nash equilibrium include quantal response equilibrium (Goeree et al., 2002) and level-k models of bidders (Crawford and Iriberri, 2007).

The purpose of this paper is not to determine which, if any, of these models best explains our data. However, it is worth noting that risk aversion, a joy of winning, or regret aversion are not able to explain both the over-entry and under-entry we observe.

Our results on bidding are surprising, and suggest a need for further research. While we are not able to conclusively offer an explanation for this behavior, We conjecture that a portion of bidders are not treating the cost of entry as a sunk cost when formulating their bids. Bidding relative to a naive model of bidding, in which a bidders behaves as though his value were $v_{it} - c_{it}$ is illustrated in Figure 5 which contains kernel densities of bids relative to this model by group size and auction format. Notice that the densities are bimodal, and that in English clock auctions, one of these modes corresponds with this naive model of bidding. Note that such a model is not able to fully explain the underbidding we observe in first-price auctions.

Another factor that may contribute to the observed underbidding in first-price auctions is the relatively high level of feedback utilized in our design. In particular, subjects are informed of all observed bids at the end of each period. Neugebauer and Selten (2006) and Isaac and Walker (1985) both find that providing feedback which includes the losing bids tends to reduce bids. Further research is needed to determine the extent to which feedback drives bidding behavior in our experiment.

[27]This need not be true. Strictly speaking, all that need hold is for the bids that determine prices to be, on average, lower than Nash predictions.

[28]Recall that the cutoff value is given by $v_c = 100 \cdot (c_i/100)^{1/n}$, so that the intercept is increasing in the realized participation cost. This is because a higher participation cost implies a higher entry threshold value, which means that in equilibrium any bidder must have a higher value.

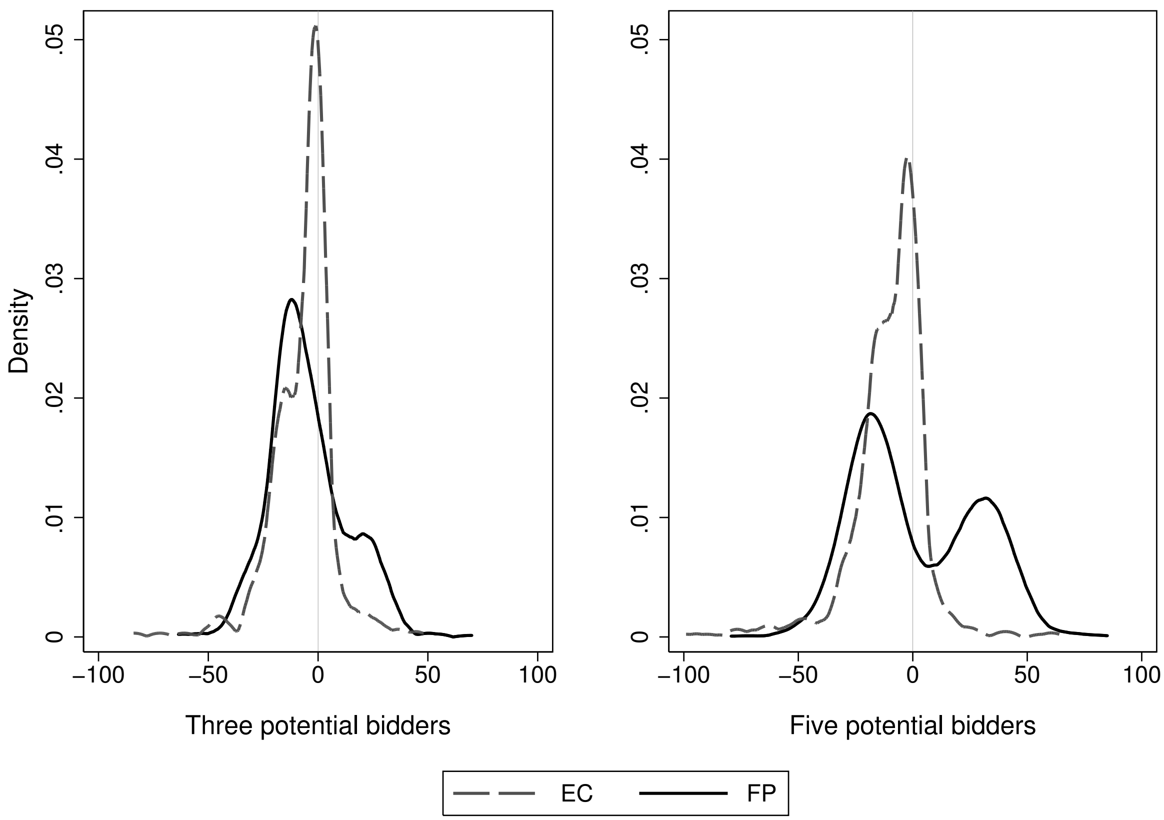

Figure 4: Kernel densities of bid deviations from Nash predictions by auction format and number of potential bidders.

Table 11: Summary statistics for bidding conditional on observed entry behavior

Treatment	Observed bids	Predicted bids	Predicted bids with sunk cost fallacy
FP3	42.888	47.647	43.168
	(18.829)	(27.703)	(24.797)
FP5	45.762	44.399	39.880
	(18.427)	(38.234)	(34.268)
EC3	45.237	51.785	34.788
	(20.623)	(23.360)	(28.611)
EC5	48.496	58.382	30.082
	(21.675)	(23.904)	(34.455)

Notes: Table contains means with standard deviations in parentheses. First-price auctions include all bidders, while English clock auctions include all non-winning bidders. Predicted bids are calculated using the realized values and entry costs utilized in the experiment.

To further analyze bidding behavior, we estimate bidding functions for each auction format via GLS and include random effects to control for individual subject variation, and cluster standard errors at the session

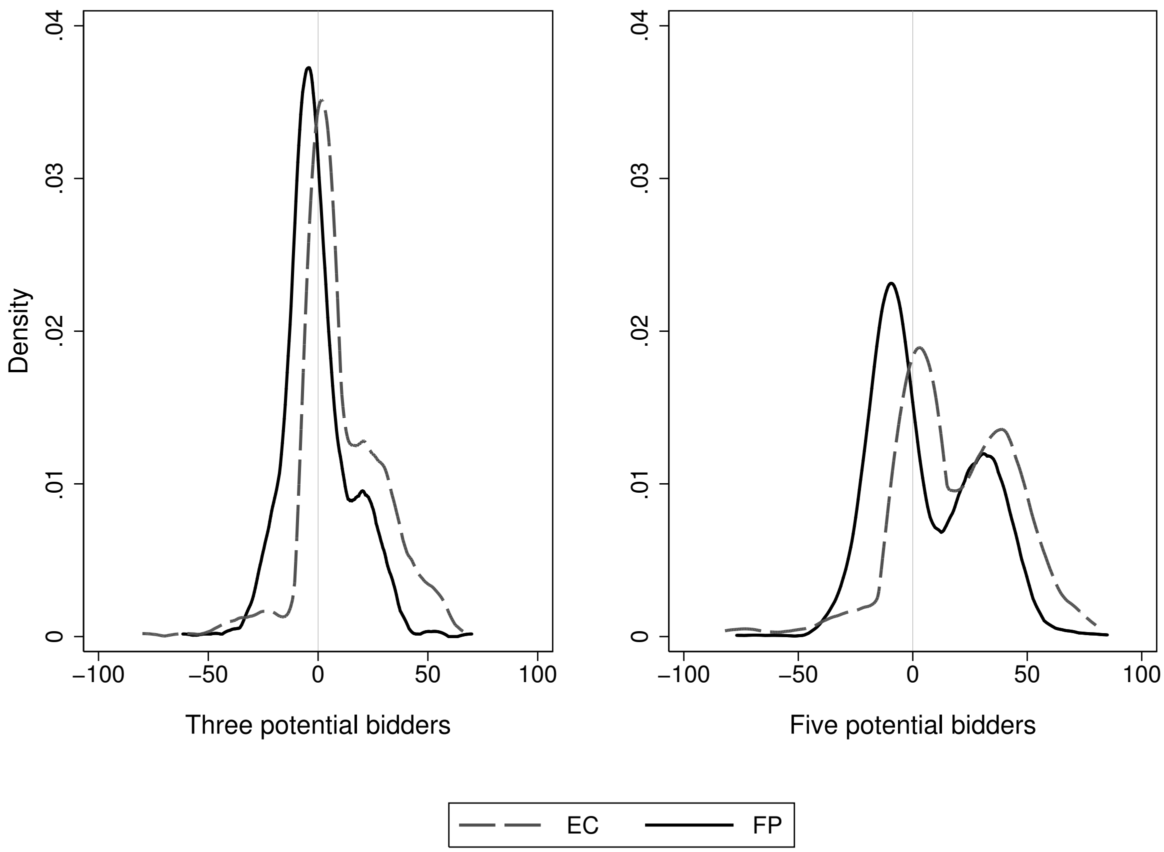

Figure 5: Kernel densities of bid deviations from naive predictions by auction format and number of potential bidders.

level. The dependent variable is the observed bid.[29] To determine the effect of value on bids, we include v_{it}. We also include the observed number of bidders (m_{it} and the realized entry cost (c_{it}) of bidder i in period t. Additionally, we include G^5_{it}, and interact this dummy with v_{it} m_{it} and c_{it}. In some specifications we also include additional controls. In particular we control for $Male_i$, Age_i, $ln(t+1)$), and $GroupOrder_i$.

Table 12 reports results for English clock and first-auctions in which there in more than one bidder.[30] Notice that in English clock auctions, the coefficient on v_{it} is predicted to be one, and all other coefficients are predicted to be zero. However, while the coefficient on v_{it} is positive and highly significant it is statistically lower than one in all specifications.[31] Thus, bidders in English clock auctions respond positively to value, but despite it being a weakly dominated bidding strategy, they bid less than their value. Note that bidders do seem to be learning, as evidenced by an increase of the coefficient for v_{it} during the second half of the experiment. However, the coefficient is still less than one. Note that he conjecture that bidders are falling victim to the sunk cost fallacy is consistent with the negative and statistically significant coefficient on c_{it}. This coefficient decreases during the second half, but it is still negative and marginally significant.

Also counter to theory, we find that under some specifications bids are increasing in m_{it} when group

[29]In English clock auctions we only observe the bid of non-winning bidders. Thus, our analysis of English clock auctions will restrict attention to non-winning bids.

[30]We restrict attention to auctions with more than one bidder because in first-price auctions with only one bidder they can win the auction with a bid of zero, and in English clock auctions with one bidder the auction ends automatically at a price of zero.

[31]The p-values for tests of coefficients are reported below the relevant specification.

Table 12: Random effects estimates of the determinants of bids in auctions with more than one bidder

	English clock			First-price		
	All 40 periods		Last 20 periods	All 40 periods		Last 20 periods
	(1)	(2)	(3)	(4)	(5)	(6)
G^5_{it}	-0.001	0.966	7.848	1.604	1.485	1.207
	(3.598)	(3.637)	(4.124)	(1.471)	(1.793)	(3.136)
v_{it}	0.747***	0.747***	0.824***	0.675***	0.671***	0.734***
	(0.040)	(0.042)	(0.033)	(0.050)	(0.050)	(0.051)
$v_{it} \cdot G^5_{it}$	-0.033	-0.042	-0.142*	-0.002	0.000	-0.009
	(0.063)	(0.061)	(0.065)	(0.022)	(0.022)	(0.050)
m_{it}	0.61	1.396***	2.036***	0.759	0.327	-0.268
	(0.513)	(0.247)	(0.390)	(0.596)	(0.269)	(0.848)
$m_{it} \cdot G^5_{it}$	-0.159	-1.579***	-3.551***	0.077	1.534	2.532
	(0.364)	(0.402)	(0.741)	(0.182)	(1.018)	(1.431)
c_{it}	-0.353***	-0.218***	-0.184*	-0.416***	-0.398***	-0.528***
	(0.045)	(0.064)	(0.094)	(0.097)	(0.091)	(0.128)
$c_{it} \cdot G^5_{it}$	0.037	-0.133	-0.183	-0.039	0.021	0.012
	(0.109)	(0.117)	(0.244)	(0.033)	(0.083)	(0.108)
$ln(t+1)$		0.753	0.367		2.464**	4.020*
		(1.176)	(6.049)		(0.898)	(2.006)
$GroupOrder_i$		7.998***	16.607**		-4.783	-9.111*
		(1.641)	(5.770)		(4.143)	(4.594)
$Male_i$		2.53	5.159		-0.104	-0.057
		(2.197)	(3.281)		(1.386)	(1.079)
Age_i		0.037	-0.547		0.333	0.218
		(0.460)	(0.686)		(0.437)	(0.467)
$Constant$	9.053***	-0.507	2.759	2.388	-10.264	-12.725
	(1.466)	(12.268)	(27.430)	(2.035)	(10.230)	(12.747)
Tests of coefficients[a]						
$v_{it} = 1$	0.000	0.000	0.000	0.000	0.000	0.000
$v_{it} + v_{it} \cdot G^5_{it} = 1$	0.000	0.000	0.000	0.000	0.000	0.000
$m_{it} + m_{it} \cdot G^5_{it} = 0$	0.417	0.336	0.047			
$c_{it} + c_{it} \cdot G^5_{it} = 0$	0.000	0.000	0.127			
Observations	837	837	407	1,428	1,428	682
Clusters	5	5	5	5	5	5
Overall R^2	0.613	0.619	0.611	0.712	0.721	0.789

Notes: Determinants of bidding using individual level random effects estimated by generalized least squares. Dependent variable is observed bid, conditional on entry. Standard errors clustered at the session level are reported in parentheses.
$^+p < 0.10, ^* p < 0.05, ^{**} p < 0.01, ^{***} p < 0.001$
$^a p$-values for each relevant specification are reported in the corresponding column.

size is three, but that this effect is negative for a group size of five. The magnitude of these effects are small and may reflect (anti) social preferences.[32]

For first price auctions, we find a positive and statistically significant effect of value, and a negative and statistically significant effect of participation cost. Both are relevant variables for Nash bidding. However, the predicated coefficients depend on the number of bidders. To facilitate the comparison on bidding behavior with Nash predictions we report additional specifications which include $\nu_{it} = v_i \cdot (m-1)/m$ (i.e. the slope of the Nash bidding function) in place of v_{it}, and $\kappa_{it} = (c_i/100)^{1/n} \cdot 100/m$ (i.e. the intercept of the Nash bidding function) in place of c_{it}.

Table 13 contain the results of these specifications. We find that the coefficient on ν_{it} is not only positive and significant, but we are unable to reject that it is equal to one in any specification. In the first half of the experiment, the interaction of ν_{it} and groups size is negative and significant, indicating that responsiveness is less than predicted by theory when group size is five.[33] However, if we restrict attention to the second half of the experiment, the coefficient on $\nu_{it} \cdot G^5_{it}$ is not longer significant: we cannot reject that bids respond to changes in value as predicted by theory, regardless of group size. Thus, the slope of our estimated bid functions are in line with theoretical predictions.

The coefficient on κ_{it} is positive and statistically significant when we consider all periods, but we reject the null that it is equal to one. Furthermore, when we restrict attention to the second half of the experiment, the coefficient is no longer significant. That is, the intercept of the bid function seems to be lower than what theory predicts. As such, we conclude that deviations of bidding from theory are largely driven by bidders not accounting for the information that the entry cost conveys regarding the interval of bids from which the values of other bidders are drawn.

[32]Cooper and Fang (2008) find evidence for spiteful bidding in second price auctions.

[33]We reject the null that the sum of the coefficients on ν_{it} and $\nu_{it} \cdot G^5_{it}$ are equal to one. p-values for these tests are reported below the relevant specifications.

Table 13: Random effects estimates of the responsiveness of bids in first-price auctions to equilibrium predictions

	All 40 periods		Last 20 periods	
	(1)	(2)	(3)	(4)
ν_{it}	1.030***	1.030***	1.089***	1.092***
	(0.071)	(0.067)	(0.084)	(0.085)
$\nu_{it} \cdot G^5_{it}$	-0.148***	-0.139***	-0.145	-0.143
	(0.036)	(0.032)	(0.076)	(0.078)
κ_{it}	0.178***	0.225**	0.134	0.178
	(0.050)	(0.075)	(0.081)	(0.093)
$\kappa it \cdot G^5_{it}$	0.220**	0.176*	0.198	0.214
	(0.085)	(0.082)	(0.132)	(0.135)
G^5_{it}	-3.18	-2.205	-3.816	-3.916
	(1.998)	(1.411)	(3.384)	(3.119)
$ln(t+1)$		3.361***		7.061*
		(0.590)		(3.398)
$GroupOrder_i$		0.021		-0.957**
		(0.639)		(0.313)
$Male_i$		-0.809		-0.343
		(1.317)		(1.057)
Age_i		0.282		0.217
		(0.421)		(0.478)
$Constant$	1.903	-14.228	2.858	-26.461
	(2.006)	(9.420)	(1.875)	(16.951)
Tests of coefficients[a]				
$\nu_{it} = 1$	0.67	0.657	0.291	0.281
$\nu_{it} + \nu_{it} \cdot G^5_{it} = 1$	0.022	0.028	0.293	0.319
$\kappa_{it} = 1$	0.000	0.000	0.000	0.000
$\kappa_{it} + \kappa_{it} \cdot G^5_{it} = 1$	0.000	0.000	0.000	0.000
Observations	1428	1428	682	682
Clusters	5	5	5	5
Overall R^2	0.67	0.688	0.729	0.735

Notes: Estimated responsiveness of first-price auction bids relative to theoretical predictions. Individual level random effects estimated via generalized least squares. Dependent variable is observed bid in first-price auction, conditional on entry and having at least one more entrant ($n > 1$). Standard errors clustered at the session level are reported in parentheses.
$^+p < 0.10, ^* p < 0.05, ^{**} p < 0.01, ^{***} p < 0.001$
[a] p-values for each relevant specification are reported in the corresponding column.

C Comparison of BDM to direct entry

As previously mentioned, all participants in the experiment which is the subject of this paper had previously participated in an experiment which also analyzed auctions with endogenous participation. The design of this experiment involved direct entry into first-price and English clock auctions, which allow us to compare behavior using BDM versus direct entry. The common opportunity cost of entry (integers between one and twenty) was common knowledge when entry decisions were made and the independent private values were drawn from a discrete uniform distribution from one to one hundred. After observing both of these pieces of information each potential bidder decided to enter the auction or not. The auction format was varied on a within subject basis, and whether or not bidders were informed of the number of entrants prior to placing their bids was varied on a between subject basis. The number of potential bidders was three. A detailed analysis of the results of this experiment can be found in Aycinena and Rentschler (2014).

To determine whether or not the use of a BDM entry mechanism yields biased results, we compare the data from this experiment with the data from Aycinena and Rentschler (2014) in which bidders knew the number of entrants when formulating bids. Since this earlier experiment always involved three potential bidders, this comparison can only be done for our treatments that also involve three potential bidders. For every possible value a potential bidder could observe we take the median revealed WTP in the relevant auction format from the experiment using the BDM entry mechanism. We then determine whether the direct entry decisions from the data of Aycinena and Rentschler (2014) is consistent with the median observed WTP.Although we present results for all periods, to ensure comparability, we will focus attention on the second half of the experiments in Aycinena and Rentschler (2014).

We have two measures of consistency. In the first, we say that a direct entry decision is consistent with the BDM entry data if either: 1) a subject entered the auction when, conditional on their value, the opportunity cost of entry was below the median revealed WTP using the BDM entry mechanism, or 2) a subject did not enter the auction when, conditional on their value, the opportunity cost of entry was higher than the median revealed WTP. The second measure of consistency distinguishes whether or not an observed entry decision that is not consistent with the median WTP under the BDM entry mechanism was due to: 1) a potential bidder in Aycinena and Rentschler (2014) entering despite the opportunity cost of entry being larger than the median WTP under the BDM entry mechanism, or 2) not entering, despite the opportunity cost of entry being lower than the median WTP under the BDM entry mechanism.

Probit estimates of the first measure of consistency of the direct entry data with the median revealed WTP are contained in Table 14. We control for auction format (FP_{it} equal to one if the subject was in a first-price auction), a subject's value (v_{it}), the common opportunity cost of the auction ($OutsideOption_{it}$). We also report specifications which control for learning ($ln(t+1)$) and the order in which a subject was exposed to the two auction formats ($AuctionOrder_i$). Lastly, we report specifications which consider all periods, as well as specifications which restrict attention to the second half of the experiment with direct entry. There are several results worth noting. Most importantly, the coefficient corresponding to auction format is insignificant at conventional levels in all specifications. Thus, the BDM entry mechanism does not seem to affect entry behavior differently across auction formats. Also, note that the average of the dependent variable is 0.77 when all periods in Aycinena and Rentschler (2014) are considered, and 0.8 when attention is restricted to the second half. This means that the BDM entry method is consistent with direct entry a large majority of the time. This is despite the fact that in Aycinena and Rentschler (2014) the entry cost is an opportunity cost (while in the current experiment it is an explicit cost), and the experience level of participants is not the same across the two data sets. Note that the coefficient corresponding to $ln(t+1)$ is positive and significant across all specifications, indicating that as subjects gain more experience their direct entry decisions are more aligned with the BDM entry decisions.

It is also important to note that a potential bidder's value is associated with a higher probability of inconsistent entry and that a higher opportunity cost of entry is associated with a lower probability of inconsistency. However the magnitude of these effects is quite small, and are smaller when attention is restricted to the second half of the data from Aycinena and Rentschler (2014). Lastly, note that the constant is positive and significant. To determine whether this is driven by the fact that we do not distinguish between the two possible types of inconsistency between the direct and BDM entry decisions, we also report ordered profits. The categorical dependent variable in these regressions is equal to negative one if entry is not predicted by the BDM but is observed in the direct entry data (direct entry is "under-predicted" by BDM), equal to zero if entry in the BDM is consistent with the direct entry data, and equal to one if entry is predicted by the BDM but is not observed in the direct entry data (direct entry is ”over-predicted” by BDM).

Results are reported in Table 15, and are largely consistent with the probit specifications. However, when attention is restricted to the second half of the data, note that the constant is no longer significant and this specification is favored by the BIC and AIC.

While this analysis is far from a systematic evaluation of the consistency of entry behavior in auctions across entry mechanisms, our results suggest that if the BDM biases behavior, there is no evidence that the bias varies by auction format.

Table 14: Probit estimates of the consistency of median WTP with direct entry

	All 40 periods		Last 20 periods	
	(1)	(2)	(3)	(4)
FP_i	-0.003	0.007	-0.014	-0.025
	(0.035)	(0.061)	(0.079)	(0.064)
v_{it}	0.007***	0.007***	0.006***	0.006***
	(0.001)	(0.001)	(0.001)	(0.001)
$OutsideOption_{it}$	-0.021***	-0.021***	-0.017***	-0.017***
	(0.003)	(0.003)	(0.005)	(0.004)
$ln(t+1)$		0.126***		-0.217
		(0.021)		(0.143)
$AuctionOrder_i$		0.040		0.109**
		(0.050)		(0.054)
$Constant$	0.675***	0.221	0.772***	1.408***
	(0.060)	(0.140)	(0.086)	(0.521)
Mean of dependent variable	0.773	0.773	0.804	0.804
Observations	5,184	5,184	2,592	2,592
Clusters	9	9	9	9
Log-likelihood	-2,714.6	-2,701.3	-1,259.1	-1,256.3
Bayesian information criterion	5,463.4	5,453.9	2,549.7	2,559.8
Akaike's information criterion	5,437.2	5,414.6	2,526.3	2,524.7

^{a}p-values for each relevant specification are reported in the corresponding column.
Standard errors in parentheses
$^+p < 0.10, ^* p < 0.05, ^{**} p < 0.01, ^{***} p < 0.001$

Table 15: Ordered probit estimates of the consistency of median WTP with direct entry

	All 40 periods		Last 20 periods	
	(1)	(2)	(3)	(4)
FP_{it}	-0.012	-0.010	0.029	0.031
	(0.019)	(0.021)	(0.031)	(0.034)
v_{it}	0.013***	0.014***	0.011***	0.011***
	(0.001)	(0.001)	(0.001)	(0.001)
$OutsideOption_{it}$	-0.006*	-0.006*	-0.011*	-0.011*
	(0.003)	(0.003)	(0.006)	(0.006)
$ln(t+1)$		0.028		0.082
		(0.024)		(0.073)
$AuctionOrder_{it}$		0.057**		0.112**
		(0.028)		(0.044)
Mean of dependent variable	0.031	0.031	0.026	0.026
Observations	5,184	5,184	2,592	2,592
Clusters	9	9	9	9
Log-likelihood	-3,371.0	-3,368.9	-1,554.9	-1,552.4
Bayesian information criterion	6,784.7	6,797.8	3,149.2	3,159.9
Akaike's information criterion	6,751.9	6,751.9	3,119.9	3,118.9

[a] p-values for each relevant specification are reported in the corresponding column.
Standard errors in parentheses
$^{+}p < 0.10, ^{*}p < 0.05, ^{**}p < 0.01, ^{***}p < 0.001$