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# Hedonic Pricing, Information, and the Market for Thoroughbred Yearlings

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Building on the 1997 work of Chezum and Wimmer, and the 1998 work of Lansford, Freeman, Topliff, and Walker, we estimated a hedonic hammer price model on a random and representative sample of 212 yearlings from the 1999 Keeneland September Yearling Sale. Explanatory variables representing day of sale, age of yearling, stud fee, racing performance of sire and dam, geographic origin of yearling, and yearling health information were statistically significant. In each model, we failed to reject the null hypothesis of no adverse selection; sellers who breed and race horses did not receive a statistically significant price penalty on their yearlings sold in this auction, compared to sellers who just breed horses.

***Key Words:*** agribusiness, equine, hedonic pricing, information, price discovery, thoroughbred yearlings

The thoroughbred industry has long been considered Kentucky's signature industry, and understandably so. In addition to its 126-year history with the Kentucky Derby, the industry epitomizes the regional culture maintaining the nation's Horse Park, several large race tracks, extensive breeding operations and horse farms, state-of-the art equine disease prevention research and veterinary hospitals, and several large thoroughbred auction houses (e.g., Lexington's Keeneland Association, Inc. is considered to be the world's leading thoroughbred auction facility).

In 1999, equine sales and stud fees totaled $830 million in gross cash receipts in Kentucky (nearly one-fourth of all agricultural cash receipts), surpassing burley tobacco to become the single largest commodity in the state's agricultural economy (Kentucky Department of Agriculture, 2000). This figure represents a 30.5% increase over the $636 million reported in 1997. Recent thoroughbred yearling auction market activity likely points to continued industry growth early into this new millennium.

Thoroughbred sellers, buyers, and industry analysts have long recognized the value of information. Leading trade publications such as *The Blood-Horse*, *The Thoroughbred Times*, *The Thoroughbred Daily News*, and *The Blood-Horse Market Watch* synthesize public auction data to provide market forecasts, analyses, and data summaries to market participants. However, it is a commonly held industry belief

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that the informational scale tips in favor of thoroughbred sellers. Breeders are typically involved with every facet of the animal's growth and preparation for the pending sale. Thus, sellers are likely aware of the animal's character, temperament, health history, and other variables which may affect future athletic ability. Buyers at the auction market often do not possess this information.

It is not uncommon for a thoroughbred buyer to hire an industry expert, such as an equine veterinarian, to serve as an agent to help reduce the risk inherent in purchasing a horse. A buyer's need for better information is underscored in yearling markets where a thoroughbred's racing ability is masked by limited physical development, but the assessment of racing conformation at this age is still an art. Buyers must rely heavily on the collective racing performance of the pedigree and other factors such as seller reputation.

Recognizing this problem, Chezum and Wimmer (1997) estimated a hedonic pricing model of final auction selling (hammer) prices to test Akerloff's (1970) Nobel Prize winning adverse selection hypothesis using a random and representative sample of 304 yearlings from the 1994 Keeneland September Yearling Sale. In the yearling auctions, some of the sellers breed horses, while other sellers not only breed, but also race. Chezum and Wimmer hypothesized sellers who breed and race receive a discount or penalty compared to the average price received by those sellers who only breed. In their hedonic pricing model, hammer prices are modeled as a function of racing intensity, pedigree of the sire and dam, age, gender, geographic origin of the foal, and other observable factors. The empirical evidence supported their hypothesis; racing intensity of the seller was statistically significant and inversely related to hammer price.

Given the difficulty, for both buyers *and* sellers alike, in determining whether or not a yearling will eventually perform well at the track, questions emerge about the adverse selection findings in the Chezum and Wimmer study. Thus, one empirical objective of this study is to revisit the adverse selection hypothesis in the thoroughbred yearling market while controlling for other observable factors (besides those examined by Chezum and Wimmer) which may affect hammer price. In particular, we use a new data set and, in our hedonic hammer price model, include previously untested explanatory variables, such as date of sale (i.e., if the yearling sold during the "select" session of the auction or not), influence of same-sired progeny, buyer visits to the on-site health record repository, advertising, mare's age, consignment size, and individual seller reputation effects.

To further help buyers and sellers of thoroughbred yearlings understand the price discovery process, a second empirical objective of this study is to extend the contribution of Lansford et al. (1998) to the case of thoroughbred yearlings. They constructed a hedonic hammer price model for race-bred yearling quarter horses consistent with this contemporary literature. Germane to their study was the calculation of "marginal values" of individual characteristics of the horses in their sample. With this information, buyers and sellers alike can potentially make better business decisions. We, too, provide these values based on our estimated hedonic pricing model for the "average" thoroughbred yearling in our sample.

## Literature Review

Price discovery research is fundamental to the study of economics and maintains a relatively long history for livestock markets. Typically, using a hedonic pricing framework, final auction prices are modeled as some function of observable animal attributes and market sale characteristics. Davis, Bobst, and Steele (1976) showed cattle weight and lot sizes are nonlinearly related to price. Kuehn (1979) determined that number of buyers, sale size, and sale type affected West Virginia feeder cattle prices. Buccola (1982) tested whether time of sale affected price. Faminow and Gum (1986) found Arizona auction market cattle prices to be a function of breed, auction location, sex-weight interactions, lot size, and weight variables.

Schroeder et al. (1988) built upon each of the previous studies to determine what factors influence feeder cattle price differentials in Kansas. Their study used data collected from seven auction markets over a seven-week and five-week period in the fall and spring, respectively. Weight, lot size and uniformity, health, horns, condition, fill, muscling, frame size, breed, and time of sale were statistically significant explanatory variables in the hedonic pricing models. However, despite controlling for a seemingly exhaustive set of observable animal and auction characteristics, no individual model explained more than three-fourths of the variation in price.

Dhuyvetter et al. (1996) and, more recently, Chvosta, Rucker, and Watts (2001) constructed similar hedonic pricing models for bulls, but included information on expected progeny differences, or EPDs. Both studies found the traditional measures of conformation did a better job of determining bull prices than did EPDs.

Research on price discovery in the equine industry is less extensive than for other livestock industries. As noted above, Chezum and Wimmer (1997) rejected the null hypothesis of no adverse selection in thoroughbred yearling markets. They found average purse, stud fee, juvenile sire, sire-mare cross, stamina, age, sex, and geographic origin of the foal statistically significant in the hedonic hammer price model.

Lansford et al. (1998) investigated race-bred yearling quarter horses and found buyers paid premiums for fillies and older yearlings. Moreover, buyers rewarded those yearlings having champion sires and dams. As previously mentioned, a hallmark of the Lansford et al. study was the calculation of "marginal values" for the individual attributes of the horses in their sample. Each "marginal value" was defined simply as the partial derivative of the estimated hedonic hammer price model with respect to the attribute of interest, evaluated at the sample means of the attributes.

Neibergs and Thalheimer (1997) estimated structural supply and demand functions over time in a recursive model. Supply variables included average prices, taxes, cost of farm production, and stud fees, while inverse demand was modeled as a function of quantity of horses, taxes, purse levels, foreign purchases, exchange rates, GDP, and number of races. They found race purses and tax policies affect thoroughbred investment decisions in a market characterized by inelastic supply and elastic demand.

Several other authors have analyzed macroeconomic conditions affecting thoroughbred yearling prices (e.g., Karungu, Reed, and Tvedt, 1993; Buzby and Jessup, 1994). Macroeconomic factors include taxes, exchange rates, interest rates, and foreign purchases. The Buzby and Jessup study was a hybrid of sorts, as it also included critical pedigree and conformation characteristics such as stud fees, racing history of the sire and dam, age, geographic origin of the foal, and gender.

Other authors have investigated the demand for horse racing that indirectly affects thoroughbred auction markets. Demand for horse racing, and hence the derived demand for upstream thoroughbred breeding, has been extensively analyzed (Thalheimer and Ali, 1992, 1995a, b, c; Ali and Thalheimer, 1997). These studies empirically analyzed the relationship between horse race wagering and telephone betting, intertrack wagering, and transportation costs.

## Data Description

For this study, we drew a random and representative sample of 212 yearlings from a population of 4,090 yearlings in the 1999 Keeneland September Yearling Sale. The scale of this auction made this sale the world's largest single offering of thoroughbred yearlings in 1999. Every price group was represented, especially the high-value upper market (yearlings valued from $100,000 and upward) which is conspicuously absent at many regional yearling auctions. Respectively, the average and median hammer prices for the sample are $77,140 and $25,500; for the population, these figures are $77,357 and $30,000. Statistically, several tests (parametric and nonparametric) confirmed the sample well represented the population with respect to hammer price, stud fee, age, and gender. (These extensive test results are available upon request from the authors.) Table 1 catalogs descriptive statistics of the variables in the hedonic hammer price model.

The Keeneland Association, Inc. computer database is the primary source of data for this model. This database contains final hammer prices, consignor, sex, age, geographic origin of the foal, and animal health and conformation information maintained in the publicly available repository at the auction site. Computer databases maintained by the Jockey Club Information Services provide detailed information regarding thoroughbred breeders, owners, and racers. *The Thoroughbred Times Buyer's Guide* (Thoroughbred Times Company, Inc., 1999) tracks the breeder (owner of the dam at time of birth) and pedigree information for each yearling in the sale. Important pedigree information includes the sire's stud fee, the dam's success as a racehorse and as a broodmare, and a statistical description of the results of similar genetic matches.

Several of the variables germane to this study include information on visits to the health record repository, presale advertising, and breeder reputation effects. These variables were quite tedious to obtain and construct. Thus, a 5% sample of the population was feasible to collect, consistent with other sample sizes in the literature, and was more than adequate for estimation and hypothesis testing. McCloskey and Ziliak (1996) have been critical of empirical studies in which sample sizes are not as carefully chosen and justified.

**Model Development**

Using the cross-sectional sample of 212 yearlings, we estimate a hedonic model of final auction hammer prices given in stylized form by:

$$\text{(1)} \quad \begin{aligned} HPRICE = {} & \beta_0 - \beta_1 RINTENS + \beta_2 SELECT - \beta_3 AGE + \beta_4 COLT \\ & - \beta_5 DI - \beta_6 CD + \beta_7 DERBY + \beta_8 SFEE + \beta_9 SIRE1 \\ & + \beta_{10} RI + \beta_{11} APRS - \beta_{12} MAGE + \beta_{13} SDX \\ & + \beta_{14} KYFOAL - \beta_{15} PROGREP + \beta_{16} REPVIS \\ & + \beta_{17} ADVERT + \beta_{18} CONSIZ \\ & + \sum_{j=1}^{15} \beta_{18+j} REPS_j + \text{error}, \end{aligned}$$

where the a priori expected qualitative relationship between each variable and hammer price (*HPRICE*) is so indicated.

Following Chezum and Wimmer (1997), we also test the adverse selection hypothesis using several racing intensity measures. Consistent with their study, we define racing intensity (*RINTENS*) to be [racing starts/(breeding starts + 1)], where breeding starts (*BSTARTS*) measures the "number of races started by horses bred by the yearling's breeder" in the previous year, and racing starts (*RSTARTS*) measures "the number of times that a yearling's breeder was the owner of a horse that started a race" in the previous year (Chezum and Wimmer, 1997, p. 523).

The publication, *The American Racing Manual*, used by Chezum and Wimmer to construct their *Racing Intensity* measure, was discontinued after 1994. We instead use the Jockey Club Information Services database. This source has the added advantage of cataloging the track performance of all horses, whereas *The American Racing Manual* reported only those breeders' (racers') horses that earned in excess of $50,000 ($30,000). In *The American Racing Manual*, any earnings below this arbitrarily chosen threshold were assumed to be zero.

The *RINTENS* variable is continuous; if adverse selection is present in the sample (i.e., reject the null hypothesis of no adverse selection and find $\beta_1 < 0$), sellers who breed and race are expected to receive a price penalty compared to sellers who only breed horses, ceteris paribus. As the intensity of the racing operation increases, so does the penalty. In other words, under adverse selection, sellers who race horses likely keep their best athletic prospects and sell lesser prospects. Buyers, anticipating this, bid less for those horses.

Following Chezum and Wimmer, we also test the adverse selection hypothesis using $\ln(RINTENS + 1)$ and $\ln(RSTARTS + BSTARTS + 1) - \ln(BSTARTS + 1)$. The latter construction of the hypothesis controls for any scale effects present in the data. In accordance with Genesove (1993), we also test for adverse selection by replacing *RINTENS* in (1) with *ADVSEL*. Seller type is not characterized as a continuum, but as two distinct types: sellers who only breed and those who breed and race. The

**Table 1. Definitions and Descriptive Statistics of Variables in Hedonic Hammer Price Models (N = 212)**

| Variable | Definition | Mean |
|---|---|---|
| *HPRICE* | Final auction hammer price ($), includes those occurrences of reserve price not attained | 77,139.623 |
| *RSTARTS* | Racing starts, for the yearling's breeder, is the number of times a horse owned by the breeder started a race in the previous year | 26.222 |
| *BSTARTS* | Breeding starts, for the yearling's breeder, is the number of races started in the previous year by the horses he/she bred | 71.825 |
| *RINTENS* | Racing intensity, equals [*RSTARTS*/(*BSTARTS* + 1)] | 0.836 |
| *SELECT* | =1 if the yearling is sold during the first four days (the "select" session) of the 11-day auction, and 0 otherwise | 0.325 |
| *AGE* | Day of year the yearling was born, ranging from 1 to 365 | 81.660 |
| *COLT* | =1 if the yearling is male, and 0 otherwise | 0.594 |
| *DI* | *Thoroughbred Times Buyer's Guide* dosage index for the yearling | 3.430 |
| *CD* | *Thoroughbred Times Buyer's Guide* center of distribution for the yearling | 0.716 |
| *DERBY* | =1 if *DI* < 4.00 and *CD* < 1.25, and 0 otherwise | 0.807 |
| *SFEE* | 1999 stud fee ($) | 21,677.679 |
| *SIRE*1 | =1 if the yearling is among the first progeny of a given sire, and 0 otherwise | 0.203 |
| *RI* | An index of the number of starts and value of earnings per start for a yearling's dam | 2.433 |
| *APRS* | For the yearling's siblings, the average purse or earnings ($) per start for a given dam | 25,820.563 |
| *MAGE* | Age of the yearling's dam in years | 11.358 |
| *SDX* | Sire-dam cross-index, increases by 1 for every stakes winning result of horses with similar genetics | 1.807 |
| *KYFOAL* | =1 for Kentucky-foaled yearlings, and 0 otherwise | 0.844 |
| *PROGREP* | Number of same-sired progeny represented within the auction | 23.387 |
| *REPVIS* | Number of visits to the public repository of health records to inspect the yearling's file | 3.651 |
| *ADVERT* | Number of advertising pages devoted to the yearling in leading industry publications | 1.753 |
| *CONSIZ* | Number of yearlings in the respective sale consignment | 124.519 |

dummy variable *ADVSEL* equals 1 if the seller bred and raced in 1998, and 0 otherwise. If adverse selection is present in the sample data, we expect *ADVSEL* to be inversely related to *HPRICE*.

Sale placement is an important consideration at major yearling markets. By design, well-bred individuals are sequenced early in the sale. The goal of this ordering process is a downward progression of the daily average hammer price through

**Table 1. Extended**

| Variable | Std. Dev. | Minimum | Maximum | Sum |
|---|---|---|---|---|
| *HPRICE* | 146,458.536 | 600.000 | 1,450,000.000 | 16,353,600.000 |
| *RSTARTS* | 92.995 | 0.000 | 715.000 | 5,559.000 |
| *BSTARTS* | 181.261 | 0.000 | 1,288.000 | 15,227.000 |
| *RINTENS* | 3.991 | 0.000 | 45.000 | 177.282 |
| *SELECT* | 0.470 | 0.000 | 1.000 | 69.000 |
| *AGE* | 33.999 | 7.000 | 153.000 | 17,312.000 |
| *COLT* | 0.492 | 0.000 | 1.000 | 126.000 |
| *DI* | 6.954 | 0.850 | 99.990 | 727.210 |
| *CD* | 0.322 | 0.000 | 1.750 | 151.770 |
| *DERBY* | 0.396 | 0.000 | 1.000 | 171.000 |
| *SFEE* | 26,647.504 | 1,000.000 | 150,000.000 | 4,595,668.000 |
| *SIRE*1 | 0.403 | 0.000 | 1.000 | 43.000 |
| *RI* | 4.500 | 0.000 | 37.670 | 515.890 |
| *APRS* | 56,556.861 | 0.000 | 525,932.000 | 5,473,959.360 |
| *MAGE* | 4.511 | 5.000 | 26.000 | 2,408.000 |
| *SDX* | 10.737 | 0.000 | 100.000 | 383.000 |
| *KYFOAL* | 0.363 | 0.000 | 1.000 | 179.000 |
| *PROGREP* | 13.228 | 1.000 | 54.000 | 4,958.000 |
| *REPVIS* | 3.563 | 0.000 | 16.000 | 774.000 |
| *ADVERT* | 2.007 | 0.000 | 8.000 | 371.740 |
| *CONSIZ* | 151.580 | 1.000 | 441.000 | 26,398.000 |

the end of the 11-day sale. The purpose is to group like individuals within the sale in order to assist buyers. Keeneland, in its administration of the sale, considers a variety of quality characteristics in assessing sale order. For an individual yearling, the quality and accomplishments of its immediate and extended family and conformation are important in designing the earliest days of the auction. Keeneland assigns those yearlings with the relatively best breeding and finest conformation to the first four sessions of the September sale, commonly referred to as the "select" or "preferred" sessions. It is frequently hypothesized that a yearling's placement within these days signals a certain level of quality to the market.

The dummy variable *SELECT* equals 1 when a yearling in the sample sold in the first four sessions, and 0 otherwise. A priori, we expect $\beta_2 > 0$. It is noted that Buzby and Jessup (1994) used auction data exclusively taken from the Keeneland July Yearling Sale; the July sale is a two-day auction and is classified entirely as "select." Chezum and Wimmer (1997) used 1994 September sales data, but did not control for "select" and "non-select" sale days even though the two distinct sessions were part of their data set.

The age and sex of a thoroughbred yearling are important factors in its valuation. All thoroughbreds born in the same calendar year, regardless of actual birth date, are assigned a common birthday of January 1. Under this method, all animals of a given crop are the same age at the sale and later at the racetrack. However, actual January-born foals may have an inherent advantage over later-born foals. Typically, earlier foals are more mature than late foals, and thus could potentially command a premium at auction.

Consistent with Buzby and Jessup (1994), we assign age values by the day of the year the associated yearling was born. For example, a January 1 birth date equals 1, and a February 1 birth date equals 32. As the units increase for the variable *AGE*, we expect *HPRICE* to decrease. Colts are usually more desirable at a yearling auction than fillies. This arises from the increased earning capacity of a colt at the racetrack and, to some extent, residual breeding value of the best colts. The gender variable, *COLT*, equals 1 for a colt and 0 for a filly. A priori, we expect *HPRICE* to be positively related to *COLT*, as was the case in Chezum and Wimmer.

Dosage index (*DI*) and center of distribution (*CD*), each a quantitative measure of pedigree, accompany every observation in the *The Thoroughbred Times Buyer's Guide*. Thoroughbred buyers and sellers closely follow each index, despite questions of its statistical merit. Each is intended to identify the influence of successful sires in the previous four generations of a given pedigree. The *DI* and *CD* estimate the existence of speed versus stamina in a pedigree. As these measures increase, the pedigree inherently favors more speed at the expense of stamina. Therefore, a horse with a high *DI* or *CD* is unlikely to have inherited the necessary stamina to capture longer distances such as one mile and one-quarter, like the Kentucky Derby. Consistent with Chezum and Wimmer, a dummy variable *DERBY* is constructed to identify those horses likely to perform well in longer races. *DERBY* is set to 1 if *DI* is less than 4.00 and *CD* is less than 1.25, and 0 otherwise. As supported by the literature, we expect *HPRICE* to decrease as *DI* and *CD* each increase. Also, we expect $\beta_7 > 0$.

A yearling's sire is another important facet of quality. The 1999 stud fee (*SFEE*) serves as a proxy for sire quality. A sire's stud fee reflects current market estimation of that stallion's ability to produce future successful racehorses. We expect a positive relationship between stud fee and hammer price. Juvenile-sired yearlings (*SIRE*1), first-crop progeny of a given stallion, possess a unique quality. First-crop sires have no previous progeny to which current yearlings may be compared. Thus, market valuation of these yearlings relies upon a relatively limited information set. The parameter estimate $\beta_9$ must consequently be determined empirically.

The racing index (*RI*) and average purse (*APRS*) estimate the dam's contributions to a yearling's pedigree. *RI* is a function of the number of starts and value of earnings per start for a yearling's dam. Successful racehorses are most likely to produce successful progeny. Therefore, we expect *HPRICE* to be positively related to *RI*. *APRS* measures the progeny's (siblings to the yearling in question) average earnings per start for a given dam. Increasing *APRS* values suggest families with greater racetrack success. Similar to *RI*, *APRS* is likely to be positively correlated with *HPRICE*. Previously untested in the literature, mare age (*MAGE*), measured in years, may also influence hammer price. If older mares produce poor runners, then *HPRICE* should be inversely related to *MAGE*.

One final pedigree variable, sire-dam cross (*SDX*), attempts to hold constant the quality of the genetic cross between the sire and dam for each yearling. Various thoroughbred families, when interbred, have histories of producing successful runners or very poor runners, depending on the genetic match. Genetic matches that have previously produced stakes winners should add market value to a similarly bred yearling. Likewise, a genetic cross with only a short history of stakes winning progeny may present a riskier purchase on the market. *SDX* increases by 1 for every stakes winning result of similar genetic matches. We expect *SDX* to be positively related to *HPRICE*, as was found to be the case by Chezum and Wimmer.

As stated earlier, the thoroughbred industry is Kentucky's signature industry, and Kentucky boasts more thoroughbred sales revenue than any other state in the nation. A popular belief among many breeders is that the area produces superior runners. We construct a dummy variable, *KYFOAL*, to track the state of origin for a yearling. This variable equals 1 for Kentucky-foaled yearlings, and 0 otherwise. We expect, a priori, *KYFOAL* to be positively related to *HPRICE*.

Product cannibalism arises when one product gains sales at the expense of a substitute product of the same producer. Same-sired yearlings being sold in the same auction may take sales from each other, so we coin the term *progeny cannibalism* to characterize this situation. To quantify this effect, we construct the variable *PROGREP* to measure the number of same-sired competitors within the auction. A priori, we expect this variable to be inversely related to hammer price.

Since July 1996, Keeneland Association, Inc. has maintained a public repository of health records for the horses being sold at auction. Prior to that time, buyers could have a horse medically evaluated by a veterinarian to better assess racing conformation. Because data from these medical evaluations were not systematically recorded, veterinarian input was not included by Chezum and Wimmer, although the buyers in their study could have had access to medical evaluations. The Keeneland repository maintains examination records, such as visual, radiographic, endoscopic, or other inspections. Many yearling markets now administer a health records repository. Yearling consignors are encouraged (but not required) to provide to the repository complete health information on each horse. Potential buyers may then inspect this information as necessary. Typically, buyers will first inspect the animal visually and then visit the repository to satisfy remaining curiosities. Repository visits per animal (*REPVIS*) will serve as a proxy for yearling quality. This variable is likely to be positively related to hammer price, as buyers are willing to expend more effort on collecting information for higher quality horses.

Reputation effects also may influence *HPRICE*. Consignors with larger consignments may use more advertising, have greater name recognition, and maintain a longer history of operation upon which buyers may gauge a consignor's reputation. We control for these potential reputation effects with several variables including advertising (*ADVERT*), consignor size (*CONSIZ*), and individual consignor reputation effects ($REPS_j$, where $j$ is the reputation effect of the $j$th consignor).

Many consignors place emphasis on the importance of advertising their consignment in the weeks leading up to a sale. *ADVERT* measures the number of advertising pages devoted to the respective consignments as they appear in the two leading industry publications, *The Thoroughbred Times* and *The Blood-Horse*, for two months prior to and one month during the yearling market. A priori, we expect *ADVERT* to be positively related to *HPRICE*. *CONSIZ* measures the total number of yearlings in the respective sale consignment. We expect *CONSIZ* to be positively related to *HPRICE* as well. To control for reputation effects, we construct dummy variables for the top 15 consignors by absolute consignment size in the sale under study. $REPS_j$ equals 1 if the yearling sold is affiliated with the $j$th consignor, and 0 otherwise. A priori, we expect a positive relationship between *HPRICE* and $REPS_j$.

Since our sample is strictly cross-sectional in nature, our approach departs from others in the literature where macroeconomic variables were employed to explain variation in hammer prices (Karungu, Reed, and Tvedt, 1993; Buzby and Jessup, 1994). In particular, Buzby and Jessup combined the macroeconomic factors from 1980 to 1990 with a cross-sectional sample of yearlings sold in each one of those years. In our study, inclusion of such factors would result in a set of vectors linearly dependent with the intercept; thus, the parameter estimates on the macroeconomic variables would not be estimable.

## Empirical Results

### *Adverse Selection*

One empirical objective of this study is to test the adverse selection hypothesis: sellers who breed and race receive a price penalty compared to sellers who only breed horses, ceteris paribus. Table 2 contains only the adverse selection hypothesis test results from seven different hedonic hammer price models. While each model controls for the other variables listed in equation (1), just the parameter estimates for the adverse selection variables are cataloged. A more complete listing of the parameter estimates in equation (1), however, is given in table 3.

To avoid heteroskedasticity and nonnormality of the empirical residuals, *HPRICE*, *SFEE*, and *APRS* in (1) were transformed using a natural logarithm. This

**Table 2. Adverse Selection Hypothesis Test Results**

| Explanatory Variable | Hedonic Hammer Price Models | | | | | | |
|---|---|---|---|---|---|---|---|
| | Model 1 | Model 2 | Model 3 | Model 4 | Model 5 | Model 6 | Model 7 |
| *RINTENS* | −0.013 (0.014) | −0.009 (0.015) | — | — | — | — | — |
| ln(*RINTENS* + 1) | — | — | −0.098 (0.102) | −0.081 (0.124) | — | — | — |
| ln(*RSTARTS* + *BSTARTS* + 1) | — | — | — | — | −0.094[a] (0.102) | −0.051[b] (0.131) | — |
| −ln(*BSTARTS* + 1) | — | — | — | — | −0.121[a] (0.105) | −0.098[b] (0.126) | — |
| *Race Zero* | — | 0.058 (0.133) | — | 0.027 (0.153) | — | 0.093 (0.177) | — |
| *Breed Zero* | — | −0.103 (0.154) | — | −0.099 (0.155) | — | 0.032 (0.233) | — |
| *ADVSEL* | — | — | — | — | — | — | −0.042 (0.117) |
| $R^2$ | 0.716 | 0.717 | 0.716 | 0.717 | 0.717 | 0.718 | 0.715 |

*Notes:* Dependent variable = ln(*HPRICE*), natural logarithm of final auction hammer price. Numbers in parentheses are standard errors. Single, double, and triple asterisks (*) denote significance at the 10%, 5%, and 1% levels, respectively.

[a] Failure to reject the null hypothesis ($p > 0.10$) of equality of parameter estimates (i.e., no scale effects) with an $F$-statistic of 0.72 and [1, 177] degrees of freedom.

[b] Failure to reject the null hypothesis ($p > 0.10$) of equality of parameter estimates (i.e., no scale effects) with an $F$-statistic of 0.57 and [1, 175] degrees of freedom.

specification is entirely consistent with Halvorsen and Pollakowski (1981), although McCloskey and Ziliak (1996) caution economists against relying too heavily on statistical considerations in regression model development. In each model, we failed to reject the null hypothesis of homoskedasticity with the Goldfeld-Quandt test and the null hypothesis of normality with the Jarque-Bera and Shapiro-Wilk tests. (These extensive test results are available upon request from the authors.)

The first six models in table 2 closely follow Chezum and Wimmer, and utilize various continuous measures of racing intensity. In models 1, 3, and 5 from table 2, we test the adverse selection hypothesis, respectively using *RINTENS*, ln(*RINTENS* + 1), and ln(*RSTARTS* + *BSTARTS* + 1) − ln(*BSTARTS* + 1). In models 2, 4, and 6, two additional variables, *Race Zero* and *Breed Zero*, are included with the foregoing intensity measures. *Race Zero*, a dummy variable, equals 1 when *RSTARTS* equals 0, whereas *Breed Zero* equals 1 when *BSTARTS* equals 0. A priori, the parameter estimate on each variable is expected to be negative. The final column of table 2 (model 7) follows Genesove's (1993) dummy variable approach in which seller type (i.e., sellers who only breed, and those who breed and race) is characterized by the variable *ADVSEL*.

Without exception, in each of the seven scenarios, we fail to reject the null hypothesis of no adverse selection at the 10% level of significance ($p > 0.10$). Our parameter estimate for *RINTENS* is −0.013, slightly smaller in magnitude than the −0.0082 obtained by Chezum and Wimmer. Statistically, however, our estimate is no different than zero. Including *Race Zero* and *Breed Zero*, each statistically insignificant ($p > 0.10$), results in a smaller estimate for *RINTENS* of −0.009, compared to −0.0075 reported by Chezum and Wimmer. Again, in contrast to their study, our parameter estimate is statistically no different than zero. Using the variable ln(*RINTENS* + 1) to test the adverse selection hypothesis yields the same results as *RINTENS*. The magnitude of our parameter estimates of −0.098 and −0.081 (models 3 and 4, table 2) are similar to those in Chezum and Wimmer (−0.1026 and −0.1404, respectively). Again, in our model, each parameter estimate is statistically insignificant ($p > 0.10$).

Testing for adverse selection with ln(*RSTARTS*+*BSTARTS* + 1) − ln(*BSTARTS* + 1) also permits the test of the null hypothesis of no scale effects. If the parameter estimate for ln(*RSTARTS* + *BSTARTS* + 1) equals the parameter estimate for −ln(*BSTARTS* + 1), no scale effects exist. In columns for models 5 and 6, we fail to reject the null hypothesis of no adverse selection ($p > 0.10$). In model 5, the magnitude of the coefficients for ln(*RSTARTS* + *BSTARTS* + 1) and −ln(*BSTARTS* + 1) of −0.094 and −0.121, respectively, parallel those in Chezum and Wimmer (−0.0638 and −0.1200).

Including *Race Zero* and *Breed Zero*, each statistically insignificant ($p > 0.10$), serves to increase the parameter estimates for ln(*RSTARTS* + *BSTARTS* + 1) and −ln(*BSTARTS* + 1) to −0.051 and −0.098. Again, both parameter estimates statistically are no different than zero. The *F*-statistic of 0.72 and [1, 177] degrees of freedom, associated with the test of no scale effects in model 5, is not statistically significant ($p > 0.10$). Similarly, the *F*-statistic of 0.57 and [1, 175] degrees of freedom, associated with the test of no scale effects in model 6, is also not statistically significant ($p > 0.10$). Chezum and Wimmer reported mixed results; in one model they found a statistically significant scale effect, but did not in the second.

As a final test of adverse selection, we used an approach similar to Genesove (1993) and classified sellers into two distinct types with the dummy variable *ADVSEL*; sellers who breed and race are set to 1, while sellers who only breed thoroughbreds are set to 0. If adverse selection was present in the sample, a priori, we expected there to be an inverse relationship between *HPRICE* and *ADVSEL*. The parameter estimate of −0.042 in table 2, however, is statistically no different than zero ($p > 0.10$); again, we fail to reject the null hypothesis of no adverse selection.

### *Estimates of Other Parameters and Marginal Values*

Table 3 contains the parameter estimates from equation (1); the adverse selection hypothesis test results for this model were also reported under model 1 in table 2 and discussed in the previous section. Since the parameter estimates across each of the seven models were so similar, only results from one model are reported and discussed

**Table 3. Parameter Estimates in Hedonic Hammer Price Model 1**

| Explanatory Variable | Parameter Estimate | Standard Error | Mean for Explanatory Variable [a] | Marginal Value [b] ($) |
|---|---|---|---|---|
| Intercept | 6.876*** | 0.862 | 1.000 | 25,046 |
| *RINTENS* | −0.013 | 0.014 | 0.836 | −46 |
| *SELECT* | 0.682*** | 0.161 | 0.325 | 2,486 |
| *AGE* | −0.004** | 0.002 | 81.660 | −14 |
| *COLT* | −0.111 | 0.122 | 0.594 | −404 |
| *DI* | 0.005 | 0.011 | 3.430 | 19 |
| *CD* | −0.136 | 0.247 | 0.716 | −494 |
| *DERBY* | −0.038 | 0.195 | 0.807 | −140 |
| ln(*SFEE*) | 0.272*** | 0.092 | 21,677.679 | 228 |
| *SIRE*1 | 0.147 | 0.149 | 0.203 | 536 |
| *RI* | 0.032** | 0.014 | 2.433 | 118 |
| ln(*APRS*) | −0.011 | 0.016 | 25,820.563 | −9 |
| *MAGE* | 0.017 | 0.017 | 11.358 | 63 |
| *SDX* | 0.011* | 0.005 | 1.807 | 39 |
| *KYFOAL* | 0.268* | 0.164 | 0.844 | 976 |
| *PROGREP* | −0.010* | 0.006 | 23.387 | −38 |
| *REPVIS* | 0.200*** | 0.018 | 3.651 | 728 |
| *ADVERT* | −0.067 | 0.072 | 1.753 | −243 |
| *CONSIZ* | 0.007 | 0.006 | 124.519 | 26 |

*Notes:* Dependent variable = ln(*HPRICE*), natural logarithm of final auction hammer price. Single, double, and triple asterisks (*) denote significance at the 10%, 5%, and 1% levels, respectively.

[a] Means taken from table 1.

[b] Marginal values calculated at sample means given in adjacent column.

here. However, the empirical results for the other six models are available upon request from the authors.

Also included in table 3 are the respective standard errors of the parameter estimates, the sample mean of each variable in the model, and the “marginal value” of each individual variable. As defined by Lansford et al. (1998), the marginal value is simply given by the appropriate partial derivative of the estimated hedonic hammer price model with respect to the attribute of interest, evaluated at the sample means of the attributes. The marginal value by definition describes the dollar worth of one additional unit of the attribute in question. This information is potentially valuable to both thoroughbred yearling buyers and sellers.

There is a statistically significant ($p < 0.01$) and, as expected a priori, positive relationship between *SELECT* and ln(*HPRICE*). Yearlings sold during the first four sessions, or the “select” or “preferred” sessions of the sale, commanded a price premium. This hypothesis has not been previously tested in the literature. The marginal value of being sequenced during the first four days of the 11-day auction is estimated to be \$2,486.

*AGE* is also statistically significant ($p < 0.05$) and, as expected a priori, inversely related to ln(*HPRICE*). Recall, *AGE* is constructed so that a yearling born January 1 is assigned the value of 1, whereas a yearling born February 1 is assigned the value of 32. Consistent with the findings of Buzby and Jessup (1994), Chezum and Wimmer (1997), and Lansford et al. (1998), our results show yearlings with more advanced physical development attain higher prices, ceteris paribus. The marginal value of being one day older is estimated to be \$14.

*COLT*, *DI*, *CD*, and *DERBY* are not statistically significant ($p > 0.10$). Buzby and Jessup also did not find gender of the yearling to be statistically significant, but Chezum and Wimmer reported male yearlings attained higher prices at auction. Lansford et al. found fillies to be more valuable in race-bred yearling quarter horse markets.

The parameter estimate for ln(*SFEE*) is statistically significant ($p < 0.01$) and, as expected a priori, positively related to ln(*HPRICE*). The marginal value is estimated to be \$228. This result, pedigree of the sire positively influences hammer price, is consistent with findings in the literature. Buzby and Jessup used two different stud fees in their composite model. The composite model combined yearling-specific traits with macroeconomic variables and performed better than the two individual models that kept the two sets of explanatory variables separate. *Old Stud Fee*, the stud fee paid at the time of sire, and *New Stud Fee*, the stud fee of the sire at the time of sale, were included in their model. Although *New Stud Fee* was not log-transformed in the Buzby and Jessup study, it matches the construction of *SFEE* in our study. Buzby and Jessup observed a statistically significant ($p < 0.01$) and positive relationship between *New Stud Fee* and hammer price.

Chezum and Wimmer also found higher stud fees, as measured in our study, led to higher auction prices. The final measure of just the influence of the sire's pedigree, *SIRE*1, is not statistically significant ($p > 0.10$). However, Chezum and Wimmer's corresponding measure, *Juvenile Sire*, was statistically significant and positively related to hammer price in each of their six models.

Of the several measures of the mare's pedigree, *RI* (the racing index) was statistically significant ($p < 0.05$) and, as expected a priori, positively related to ln(*HPRICE*). Recall, *RI* is a function of the number of starts and value of earnings per start for a yearling's dam; the more productive the dam is at the track, the greater is the value of her progeny. The marginal value of a one-unit increase in the racing index is estimated to be \$118. This is consistent with the findings of Lansford et al. Also, *RI* seems to be a reasonable proxy of the *Mare Standard Starts Index* used by Chezum and Wimmer. Neither ln(*APRS*) nor *MAGE* is statistically significant ($p > 0.10$). Chezum and Wimmer found *Average Purse* of the mare, equivalent to *APRS* here, to be statistically significant and positively related to hammer price. Chezum and Wimmer also included one other measure of the mare's pedigree, *First Mare*. We excluded this variable from our model given the lack of statistical significance in the Chezum and Wimmer study. *SDX* is statistically significant ($p < 0.10$) and, as expected a priori, positively related to ln(*HPRICE*). As the stakes winning results of similar genetic matches increase, so does hammer price. Chezum and Wimmer also confirmed this result in their sample of yearlings. In our model, the marginal value of a one-unit increase in the sire-dam cross is estimated to be $39.

*KYFOAL* is statistically significant ($p < 0.10$) and, as expected a priori, positively related to ln(*HPRICE*); Kentucky foals commanded higher prices than non-Kentucky foals. The marginal value of being foaled in Kentucky is estimated to be $976. This result contradicts Chezum and Wimmer, who found non-Kentucky foals received higher prices at auction. Buzby and Jessup (1994) also tested this hypothesis, but determined the variable to be statistically insignificant ($p > 0.10$). This issue appears to be unresolved in the literature and may depend on the sample drawn and the time frame in which it was drawn.

Previously untested in the literature, in addition to *MAGE* and *SELECT*, we review the empirical results for *PROGREP*, *REPVIS*, *ADVERT*, and *CONSIZ*. *PROGREP* is statistically significant ($p < 0.10$) and, as expected a priori, inversely related to ln(*HPRICE*). As the number of same-sired yearlings increases at the auction, hammer price falls, ceteris paribus. Thus, *progeny cannibalism* exists in the sample. The marginal value of a one-unit decrease in *PROGREP* is estimated to be $38.

*REPVIS* is important to this literature as it attempts to quantify the amount of information the buyer has access to regarding the health and potential racing ability of a given yearling. We argue this information may reduce information asymmetries that are believed to favor sellers in this class of markets. *REPVIS* is statistically significant ($p < 0.01$) and, as expected a priori, is positively related to ln(*HPRICE*). As buyers expend more effort to collect health information for a particular yearling, they are willing to pay more for the yearling. The marginal value of a one-unit increase in *REPVIS* is estimated to be $728.

Neither *ADVERT* nor *CONSIZ* are statistically significant ($p > 0.10$), so the variation in hammer price could not be statistically explained by advertising and consignment size. While not reported here to conserve space, the parameter estimates for the 15 seller reputation dummies are available upon request from the authors. Across the seven models, only one of the reputation dummies, for the consignor Fitzgerald/Keogh, is statistically significant ($p < 0.10$) and, contrary to a priori expectations, negative.

## Summary

In this analysis, we have estimated a hedonic hammer price model and revisited the adverse selection hypothesis in the context of thoroughbred yearling markets. Building upon the work of Chezum and Wimmer (1997), we replicated their hedonic hammer price modeling approach on a different, yet comparable, random and representative sample of 212 yearlings from the 1999 Keeneland September Yearling Sale. Controlling for pedigree of the sire and dam, age, gender, geographic origin of the foal, and several other variables, we failed to reject the null hypothesis of no adverse selection; sellers who breed and race horses do not receive a statistically significant penalty on average hammer price compared to sellers who just breed horses, ceteris paribus.

This result contrasts with the findings of Chezum and Wimmer, yet was obtained using six continuous measures of racing intensity and a dichotomous measure of seller type as well. One possible explanation for the difference in results could be due to sampling. Another explanation may be the additional variables included in our model, but not included in theirs; we investigated several previously untested variables which could possibly influence the final auction hammer price of a yearling. These include date of sale (i.e., during the "select" session of the auction or not), influence of same-sired progeny, buyer visits to the on-site health record repository, advertising, mare's age, consignment size, and individual seller reputation effects. The first three variables were statistically significant and were consistent with a priori expectations.

Our model also extended the work of Lansford et al. (1998) for the case of thoroughbred yearlings. In particular, we estimated the "marginal value" of each explanatory variable in the model to highlight the usefulness of the results to both buyers and sellers. In this framework, the marginal value by definition describes the dollar worth of one additional unit of the attribute in question.

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