Original article

Analysis of factors affecting length of competitive life of jumping horses

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Summary − Official competition data were used to study the length of competitive life in jumping horses. The trait considered was the number of years of participation in jumping. Data included 42 393 male and gelded horses born after 1968. The competitive data were recorded from 1972 to 1991. Horses still alive in 1991 had a censored record (43% of records). The survival analysis was based on Cox's proportional hazard model. The independent variables were year, age at record, level of performance in competition (these three first variables were time dependent), age at first competition, breed and a random sire effect. The prior density of the sire effect was a log gamma distribution. The maximization of the marginal likelihood of the γ parameter of the gamma density gave an estimate of the additive genetic variance. The baseline hazard, the fixed effects and the sire effects were then estimated simultaneously by maximizing their marginal posterior likelihood. Jumping horses were culled for either involuntary or voluntary reasons. The involuntary reasons included the management of the horse, for example, the earlier a horse starts competing the longer he lives. The voluntary reasons related to the jumping ability: the better a horse, the longer he lives (at a given time, an average horse is 1.6 times more likely to be culled than a good horse with a performance of one standard deviation above the mean). The heritability of functional stayability was 0.18. The difference in half-lives of the progeny of two extreme stallions exceeded 2 years.

horse / jumping / survival analysis / longevity

Résumé − Analyse des facteurs de variation de la durée de vie en compétition des chevaux de concours hippique. *La durée de vie sportive des chevaux de concours hippique est analysée à partir des données des compétitions officielles. Le caractère étudié est le nombre d'années en compétition. Les données concernent 42 393 chevaux mâles et hongres nés depuis 1968 et enregistrés en compétition de 1972 à 1991. Les chevaux encore en compétition en 1991 se voient attribuer une donnée dite censurée (43 % des données). L'analyse de survie est basée sur le modèle de risque proportionnel de Cox. Les variables indépendantes sont l'année, l'âge au moment de l'enregistrement, l'âge à la première*

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compétition, le niveau de performance en compétition, la race et un effet « père » aléatoire. La densité a priori de l'effet « père » est une distribution log gamma. La maximisation de la vraisemblance marginale du paramètre γ de la fonction de densité gamma permet une estimation de la variance génétique additive. La fonction de risque de base, les effets fixés et l'effet « père » ont été estimés de façon simultanée par la maximisation de leur vraisemblance marginale a posteriori. Les chevaux de concours hippique sont éliminés de la compétition soit pour raisons volontaires, soit pour raisons involontaires. Les premières sont dues aux circonstances (effet année) et à la valorisation : plus un cheval commence tôt la compétition, plus il y reste longtemps. Les secondes concernent l'aptitude du cheval au saut d'obstacles : meilleur est le cheval, plus longtemps il concourt (à un moment donné, un cheval moyen a 1,6 fois plus de chances d'être éliminé qu'un bon cheval de performance égale à un écart type au dessus de la moyenne). L'héritabilité de la longévité fonctionnelle est 0,18. La différence entre les demi-vies des descendants de deux étalons extrêmes dépasse 2 ans.

cheval / concours hippique / analyse de survie / longévité

INTRODUCTION

The primary trait required for a jumping horse is its ability to jump obstacles. Since this requires a long training period, involuntary culling of a horse always represents an important economic loss. The reasons for culling are various and are seldom recorded because of veterinary professional secrecy. The most frequent reasons are probably lameness and breathing diseases as well as accidents and colics. Since data on specific diseases were lacking, the aggregate trait, length of competitive life, was studied to measure physical stamina and endurance. This trait includes two different aspects. Culling may be voluntary, ie, the horse does not perform at the desired level, or involuntary, ie, the horse can no longer perform at all. Two stayability traits may be defined (Ducrocq, 1988a): the 'observed' stayability, which combines sport capacity and physical resistance, and the 'functional' stayability, which measures the robustness of the horse for a given jumping quality. It is this latter trait that will be examined in this study.

MATERIALS AND METHODS

Data

The annual results of all horses in jumping competitions in France from 1972 to 1991 were available. For each horse participating in any competition, the number of competitions it started and the money it earned were recorded. However, it was not known whether the first recorded year of a horse was really its first year in competition, nor if its last recorded year was its last year in competition. According to competition rules, jumping may start from 4 years of age and continue for an indefinite period of time. Only the year of performance was recorded, as no more accurate date was available. Different measures of the length of competitive life might be used: the difference between the first and the last year in competition, the true number of years in competition (years without a start omitted), the number of starts (in this case, the scale of time is 'one' start). The true number of years in competition was considered as the most appropriate criterion.

Only males and geldings were studied. The competitive lives of males and females are quite different and should not be compared. The careers of mares are interrupted by reproduction, whereas stallions can breed and compete in the same year. Consequently, sport longevity of females is more difficult to interpret.

A general characteristic of survival analysis is censoring. Some horses began their jumping life before the beginning of data collection (left censoring). On the other hand, at the date of analysis, a large number of horses were still in competition (right censoring). In both situations, their true length of competitive life was not known, only a lower bound was known. To avoid left censoring, data of horses born before 1968 (aged more than 4 years old in 1972 and perhaps already in competition before this time) were deleted because the estimation of the parameters of the model requires the full knowledge of the past life of the horse. They represented 10.9% of the total number of horses. For horses still in competition in 1991, 31.6% of the total, the lengths of life were treated as right censored in the analysis. The same was true for exported horses (6.4% of the number of horses) and some national stallions (0.4%), which returned to the stud after some limited participation in special jumping tests. The horses reimported during their competitive life were excluded from analysis (0.3%).

Edited data included 42 393 lengths of jumping life, out of which 43.3% were censored. This represented 155 570 years of performance.

Survival analysis and derivation of the likelihood

The basic information concerning survival analysis may be found in Kalbfleisch and Prentice (1980). Only definitions of specific functions are presented here, and the form of likelihood when censoring is present. Letting T be the random variable representing the failure time (or the length of competitive life) of a horse, the survivor function is defined by:

$$S(t) = P(T \geqslant t) = 1 - F(t) \qquad [1]$$

with $F(t)$ the cumulative distribution function. The hazard function $\lambda(t)$ is defined as the instantaneous rate of failure at time t:

$$\lambda(t) = \lim_{\delta t \to 0^+} \frac{P(t \leqslant T < t + \delta t | T \geqslant t)}{\delta t} = \frac{f(t)}{S(t)} \qquad [2]$$

where $f(t)$ is the probability density function of T.

According to the Cox model (1972), the hazard function is divided into the product of two terms: the first depends only on time and represents a type of mean, the baseline hazard explaining the common aging of horses; the second depends on the explanatory variables. For a horse i:

$$\lambda(t, \mathbf{z}_i) = \lambda_0(t) \exp(\mathbf{z}_i' \boldsymbol{\beta}) \qquad [3]$$

where $\lambda_0(t)$ is the baseline hazard function, $\mathbf{z}_i$, the design vector of explanatory variables for the horse i and $\boldsymbol{\beta}$ the vector of effects of these variables. With this model, the ratio of hazards for two horses at any time depends only on covariates.

Cox (1975) proposes a method based on a partial likelihood to estimate the parameters of the hazard function. He compares the hazard of one individual who fails at time t to the hazards of the whole population alive at time t. However, this method can not be applied here because the data are annually recorded, and many horses fail at the same time. As Cox's approach is not suited to situations with a large number of ties, the following alternative likelihood must be used (Kalbfleisch and Prentice, 1980):

$$L \propto \prod_{i=1}^{n} \lambda(t_i)^{\delta_i} S(t_i) \qquad [4]$$

where L is the likelihood of all the observations, n is the number of horses in the data file, $\delta = 0$ for censored observations and $\delta = 1$ for uncensored observations. This likelihood assumes that the censoring process is independent of the explanatory variables of the length of life. Note that it requires the horse's entire competitive life history and not only its state at the time of failure.

In the case of discrete failure times such as in the present study, the particular following of the survivor function is applied from Prentice and Gloeckler (1978). The time intervals are denoted A_j and defined by:

$$A_j = [a_{j-1}, a_j] \quad \text{with} \quad \begin{cases} j = 1, \ldots, m \\ a_0 = 0 \\ a_m = +\infty \end{cases} \qquad [5]$$

A culling or censoring event during the time interval A_j is denoted t^j. For example, a horse that disappears after 3 years of competition fails at time 3. A horse that has been competing for 3 years in 1991 (last year of recording) is censored at time 3. We have:

$$S(t^j, \mathbf{z}_i) = P(T_i \geqslant a_{j-1}) = \left[\exp\left(- \int_0^{a_{j-1}} \lambda_0(x) \mathrm{d}x \right) \right]^{\exp(\mathbf{z}_i' \boldsymbol{\beta})} \qquad [6]$$

$$S(t^j, \mathbf{z}_i) = \prod_{k=1}^{j-1} \alpha_k^{\exp(\mathbf{z}_i' \boldsymbol{\beta})} \qquad [7]$$

where $\alpha_k = \exp\left(- \int_{a_{k-1}}^{a_k} \lambda_0(x) \mathrm{d}x \right)$.

The hazard function during the time interval is similarly written as:

$$\lambda(t^j, \mathbf{z}_i) = P(a_{j-1} \leqslant T_i < a_j | T_i \geqslant a_{j-1}) \qquad [8]$$

$$\lambda(t^j, \mathbf{z}_i) = 1 - \alpha_j^{\exp(\mathbf{z}_i' \boldsymbol{\beta})} \qquad [9]$$

The likelihood is then proportional to:

$$L \propto \prod_{k=1}^{m} \left[\prod_{i \in D_k} \left(1 - \alpha_k^{\exp(\mathbf{z}_i' \boldsymbol{\beta})} \right) \prod_{i \in R_k} \alpha_k^{\exp(\mathbf{z}_i' \boldsymbol{\beta})} \right] \qquad [10]$$

where D_k is the set of horses culled and R_k the set of horses alive during the time interval k.

Model

Different models of the hazard function were used to analyze the different causes of culling and the appropriate associated covariates. Each additional covariate was included in the successive models and was tested with the likelihood ratio test. The final model was:

$$\lambda(t, \mathbf{z}_i) = \lambda_0(t)\, \exp[\mathbf{z}_{Y,i}(t)'\boldsymbol{\beta}_Y + \mathbf{z}_{A,i}(t)'\boldsymbol{\beta}_A + \mathbf{z}'_{F,i}\boldsymbol{\beta}_F + \mathbf{z}_{P,i}(t)'\widehat{\boldsymbol{\beta}}_P + \mathbf{z}'_{s,i}\mathbf{s}] \quad [11]$$

where $\mathbf{z}_i(t)$ corresponds to the time-dependent covariates. The use of time-dependent covariates modeled effects that are not constant throughout the life of a horse. For example, 'year' changed each time interval and 'level of performance', (computed annually), was not constant. We denoted:

$\boldsymbol{\beta}_Y$ is the vector of 'year' effects. It included 19 levels (from 1972 to 1990). Because the year 1991 contained only censored data, its effect was not estimable.

$\boldsymbol{\beta}_A$ is the vector of 'age' effects. Usually, this effect is described by the baseline hazard function. In the present study, the baseline hazard function described the survival process with regards to the number of years in competition. However, this number of years in competition might differ from age, because the age at which a horse first competes varies, and because the horses might have years without any performance. Hence, an accurate description of the aging effect is required to explicitly include an age factor, which was defined with 15 levels: from 4 to 18 years old and more, in steps of year.

$\boldsymbol{\beta}_F$ is the vector of 'age at the first start' effects. The baseline hazard function measured the effect common to horses with the same number of years in competition; the 'age' effect measured the effect common to horses at the same age, at different moments of their competitive life. The 'age at first start' effect would measure the influence of age at first start on the whole competitive life. This effect had six levels: from 4 to 9 years old and more, in steps of 1 year.

$\widehat{\boldsymbol{\beta}}_P$ is the vector of estimates of the 'level of performance' effects. We wanted to take into account the voluntary culling of horses for reasons of lack of quality. The major problem was to choose a measure of the level of performance for each year, which remained as independent as possible of the chance of an involuntary failure in this year. Unfortunately, all measures based on earnings, including earnings per start or earnings regressed on the number of starts, were related to the number of annual starts. In addition, the number of starts was partially related to the possibility of failure in the year: the horses culled during a year had a smaller number of starts than horses remaining alive throughout this year. To assess the influence of the level of earnings regardless of the influence of the number of starts, an auxiliary model was used. This auxiliary model was defined in order to obtain adjustment factors for earnings, as independent as possible of the number of starts. Consequently, this model included a 'number of starts' effect and a 'Log(earnings)' effect, in order to separate them.

This model was:

$$\lambda(t, \mathbf{z}_i) = \lambda_0(t) \exp[\mathbf{z}_{Y,i}(t)'\boldsymbol{\beta}_Y + \mathbf{z}_{A,i}(t)'\boldsymbol{\beta}_A + \mathbf{z}'_{F,i}\boldsymbol{\beta}_F + \mathbf{z}_{N,i}(t)'\boldsymbol{\beta}_N + \mathbf{z}_{P,i}(t)'\boldsymbol{\beta}_P + \mathbf{z}'_{s,i}\mathbf{s}]$$
$$[12]$$

where $\boldsymbol{\beta}_N$ is the vector of 'number of starts' effects and $\boldsymbol{\beta}_P$ is the vector of 'level of performance' effects. This auxilary model could not be the true model, because the correction for number of starts is the correction for the longevity itself. The model with only a 'level of performance effect' would have had the same problem. But 'earnings' effects, estimated in this auxiliary model and assumed to be independent of the number of starts, were used as preadjustment factors $(\widehat{\boldsymbol{\beta}}_P)$ in the final model [19], which did not include the effect of the number of starts.

The 'number of starts' effect had eight levels: six levels from 1 to 30 starts in steps of five starts, one level from 31 to 40 starts and one level for more than 40 starts. Because the number of starts for young horses was limited by regulation, only the first three and five levels were considered at the age of 4 and 5 years, respectively. The logarithm of earnings was standardized by age and year (mean 100, standard deviation 20), assuming that the culling choice was between horses in the same year of performance and age group. Horses aged 4 and 5 years had special competitions reserved for their age class, whereas after 6 years, a horse was compared to any other horse of any age. Consequently, the level of performance was defined within these three age classes. Nine levels of performance were defined: one for the horses that did not earn any money (30% of horses each year), six between 70 and 130 in steps of ten and two at the extremes ($\leqslant 70$ and > 130). At 4 years old, the extreme classes were merged and only seven levels were considered, because the distribution deviated too much from a normal one, and because the variance was too small.

Here, $\mathbf{s}$ is the vector of 'sire' effect. This effect was the only random effect. The horses were the offspring of 4 851 sires, each with 8.7 offspring on average. More than 800 sires had over 15 offspring. No 'breed' effect was included simultaneously with the sire effect because the breed of the sire did not determine the breed of the progeny. Another model was applied to estimate breed differences:

$$\lambda(t, \mathbf{z}_i) = \lambda_0(t) \exp[\mathbf{z}_{Y,i}(t)'\boldsymbol{\beta}_Y + \mathbf{z}_{A,i}(t)'\boldsymbol{\beta}_A + \mathbf{z}'_{F,i}\boldsymbol{\beta}_F + \mathbf{z}_{P,i}(t)'\widehat{\boldsymbol{\beta}}_P + \mathbf{z}'_{B,i}\boldsymbol{\beta}_B] \quad [13]$$

where $\boldsymbol{\beta}_B$ is the vector of 'breed' effect. Three types of breeds were detected: (1) riding horse breeds including the 'Selle Français' (SF), selected mainly for jumping and representing the majority of the jumping population (59%), the 'Anglo-Arabe' (AA), selected for multiple sports (11%) and the 'Cheval de Selle' (10%), (2) race breeds including the thoroughbred (PS) for galloping races (8%) and the 'Trotteur Français' (TF) for trotting races (9%), and (3) breeds of small size horses, including ponies and Arabs (2%). An additional class included horses of unknown origins or foreign horses (0.7%).

Prior density

The sire distribution is usually assumed to be a normal one. But, in the present model, the additive polygenic effect might be defined on the exponential scale $\exp(\mathbf{s})$ (denoted $\mathbf{w}$) or on the scale of $\mathbf{s}$. To make the distribution of $\mathbf{w}$ more flexible, a

gamma density with parameters γ and γ was chosen as a prior density, as in Ducrocq et al (1988a, b); ie:

$$p(w/\gamma) = \frac{\gamma^\gamma w^{(\gamma-1)} e^{-\gamma w}}{\Gamma(\gamma)} \qquad [14]$$

where Γ is the gamma function.

The estimate of γ gave the variance of w: $V(w) = 1/\gamma$ and of $s = \log(w)$: $V(s) = \mathbf{\Psi}^{(1)}(\gamma)$ where $\mathbf{\Psi}^{(1)}$ is the trigamma function. The expectations were $E(w)$ and $E(s) = \mathbf{\Psi}(\gamma) - \log(\gamma)$ where $\mathbf{\Psi}$ was the digamma function. Sires were assumed to be unrelated.

Estimation of parameters

The *a posteriori* density of the parameters given the data was proportional to the product of the likelihood [10] by the prior density [14]:

$$f(\boldsymbol{\beta}, \boldsymbol{\alpha}, \gamma | Y) \propto \prod_{k=1}^{m} \left[\prod_{i \in D_k} \left(1 - \alpha_k^{\exp(z_i'\boldsymbol{\beta})}\right) \prod_{i \in R_k} \alpha_k^{\exp(z_i'\boldsymbol{\beta})} \right] \prod_{l=1}^{\eta} \frac{\gamma^\gamma w_1^{(\gamma-1)} e^{-\gamma w_l}}{\Gamma(\gamma)} \qquad [15]$$

where $\boldsymbol{\beta} = (\boldsymbol{\beta}_Y, \boldsymbol{\beta}_A, \boldsymbol{\beta}_F, \boldsymbol{\beta}_P, \boldsymbol{\beta}_N, \mathbf{s}), \boldsymbol{\alpha} = (\alpha_1, ...\alpha_m)$ is the survivor function by time intervals and η is the number of sires. Let $\boldsymbol{\beta} = (\mathbf{b}, \mathbf{s})$ where $\mathbf{b} = (\boldsymbol{\beta}_Y, \boldsymbol{\beta}_A, \boldsymbol{\beta}_F, \boldsymbol{\beta}_P, \boldsymbol{\beta}_N)$.

The introduction of the different fixed effects was tested by maximization of the logarithm of the likelihood alone. Then, the marginal *a posteriori* density of $\gamma(f(\gamma))$ after integration of all the effects $\mathbf{b}, \mathbf{s}$ and $\boldsymbol{\alpha}$, was used to estimate the parameter γ. This allowed us to take into account the uncertainty of the estimates of the location parameters $\mathbf{b}, \mathbf{s}$ and $\boldsymbol{\alpha}$ in the estimation of dispersion parameters. The integration of $\mathbf{b}, \mathbf{s}$ and $\boldsymbol{\alpha}$ could not be performed algebraically. On the other hand, the uncertainty was not of the same order for all the parameters. The fixed effects and the survival by time intervals were estimated from large samples, in contrast to the sire effects. Consequently, the integration of the sire effects was more necessary than that of the other effects. So instead of $f(\gamma)$, attention was paid to the marginal likelihood $f(\mathbf{b}, \boldsymbol{\alpha}, \gamma)$. This marginal likelihood could have been calculated by numerical integration of the sires, but the numerical maximization of this function, which depended on about 100 variables, with a 'quasi-Newton' algorithm, would have required more than 20 000 evaluations of the function. Because each calculation of this function required as many integrals as sires (4 851), this maximization was considered to be impossible within a reasonable computing time. Consequently, this function was approximated by the following likelihood:

$$f(\mathbf{b}, \boldsymbol{\alpha}, \gamma | Y) \approx f(\gamma | Y, \mathbf{b} = \widehat{\mathbf{b}}, \boldsymbol{\alpha} = \widehat{\boldsymbol{\alpha}}) \qquad [16]$$

This marginal likelihood required the same integration effort but depended on only one variable and was easier to maximize, provided that good $\widehat{\mathbf{b}}$ and $\widehat{\boldsymbol{\alpha}}$ values were available. These values were obtained by the maximization of $f(\mathbf{b}, \boldsymbol{\alpha}, \mathbf{s} | Y, \gamma = \widehat{\gamma})$, with the parameter γ estimated by the maximization of the preceding marginal likelihood. This defined an iterative process: $f(\gamma | Y, \mathbf{b} = \widehat{\mathbf{b}}, \boldsymbol{\alpha} = \widehat{\boldsymbol{\alpha}})$ was maximized,

giving an estimate of γ to be used in the calculation of $f(\mathbf{b}, \boldsymbol{\alpha}, \mathbf{s}|Y, \gamma = \widehat{\gamma})$, which was maximized to obtain $\widehat{\mathbf{b}}$ and $\widehat{\boldsymbol{\alpha}}$. The estimates $\widehat{\mathbf{b}}$ and $\widehat{\boldsymbol{\alpha}}$ were used again to calculate a new function $f(\gamma|Y, \mathbf{b} = \widehat{\mathbf{b}}, \boldsymbol{\alpha} = \widehat{\boldsymbol{\alpha}})$, which was maximized to obtain a new $\widehat{\gamma}$. At convergence, the γ value was expected to be close to the one that would maximize $f(\gamma|Y)$.

The numerical integration of the sires was performed using the NAG (1991) subroutine D01BAF. The maximization of $f(\gamma|Y, \mathbf{b} = \widehat{\mathbf{b}}, \boldsymbol{\alpha} = \widehat{\boldsymbol{\alpha}})$ was obtained by the NAG (1991) subroutine E04ABF. The maximization of $f(\mathbf{b}, \boldsymbol{\alpha}, \mathbf{s}|Y, \gamma = \widehat{\gamma})$ was obtained by a Newton–Raphson algorithm. The solutions of the system were obtained by absorbing the equations corresponding to sire effects, taking advantage of the diagonal structure of the corresponding matrix of second derivatives. The final solutions for fixed effects and sire effects were obtained by maximizing $f(\mathbf{b}, \boldsymbol{\alpha}, \mathbf{s}|Y, \gamma = \widehat{\gamma})$ after convergence for γ.

RESULTS

Convergence of the algorithms

Maximizing the logarithm of the likelihood alone by a Newton–Raphson algorithm was very fast. Six iterations were usually required. The square root of the ratio of the squared difference of the logarithm of the likelihood between two iterations and the squared value of this likelihood was less than 10^{-13} and the same criterion applied to the solutions of fixed effects and sire was less than 10^{-15}. The convergence of the γ parameter of the gamma function of the a priori density of sires was also fast. The maximization algorithm found the new parameters in usually eight calls to the function. The iterations between the two functions maximized were stopped when the parameter γ was known with an accuracy of 0.01.

Choice of the model

Three causes of involuntary culling were retained from the results of table I: calendar year, age and age at first start. The interaction between age and age at first start was removed. The introduction of 'level of performance' effect, the voluntary cause of culling, greatly increased the likelihood. The parameter estimates presented below are those obtained with a sire model after convergence for γ.

Distribution of the length of jumping life

The 'α' parameters (survival in time interval), 'age' effects and 'age at first start' effects can only be combined in certain ways. Survivor function, density function and hazard function were reconstructed for each class of age at first start. For example, probability of remaining 3 years in competition for a horse that started at 5 years old was the combination of survival at 3, age 8, first start 5.

Figure 1 diplays the density function for horses differing in age at their first start. For those horses that started at younger ages (4–5 years), the curve is quite flat during the first years of competition (equal probability, 8%, of remaining 1–7 years

Table I. Likelihood ratio test on fixed effects.

	Model		$-$ 2 Deviation Log(likelihood)	df	χ^2 P < 0.001
(1)	Year	(1) $-$ KM[a]	1 068.5	18	42.3
(2)	Year + age	(2)–(1)	4 322.8	14	36.1
(3)	Year + age + age at first start	(3)–(2)	267.2	5	20.5
(4)	Year + age $\times$ age at first start	(4)–(3)	104.5	55	93.2
(5)	Year + age + age at first start + number of starts + level of performance	(5)–(3)	10 586.1	35	66.6
(6)	Year + age + age at first start + adjustment for level of performance	(6)–(3)	7 012.6	0	–
(7)	Year + age + age at first start + adjustment for level of performance + breed	(7)–(6)	180.0	6	22.5

[a] The deviation is between the log-likelihood without covariates that corresponds to classical estimation of survival of Kaplan and Meier (1958).
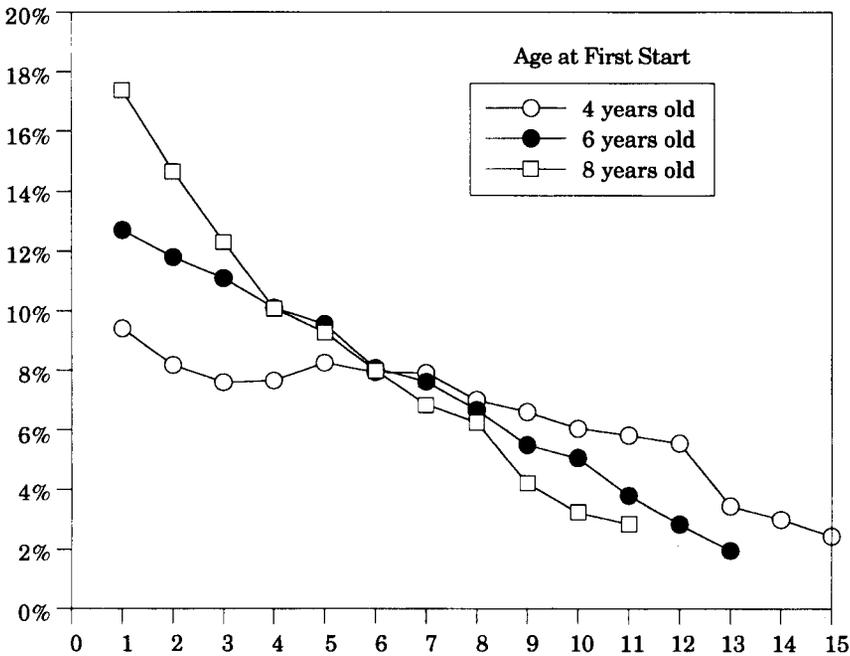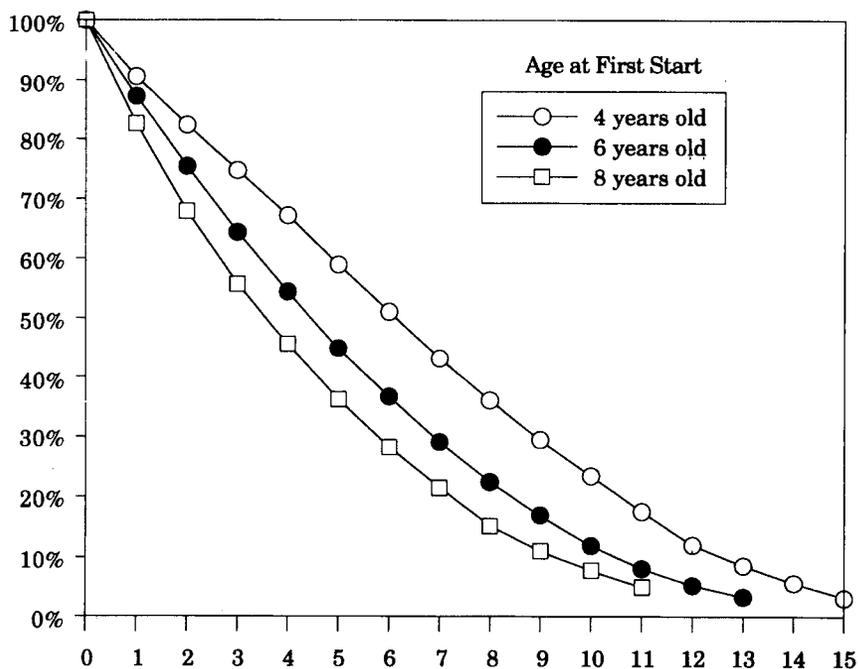

Fig 1. Density function of the length of jumping life: number of years in competition (x axis) versus probability (y axis).

in competition). In contrast, when horses began after 6 years, the density function always decreased and the slope increased with the age at first start.

The survivor function curves (fig 2) never overlapped: the probability of still competing after any number of years in competition was always greater for horses that started the competition earlier. However, the phenomenon was not strong enough for the probability of still being alive at a given age to remain higher for horses that started earlier, because the number of years in competition was higher for horses that started earlier. The probability of still remaining after 5 years in competition was 59, 53, 45 and 41%, for horses beginning at 4, 5, 6 and 7 years old, respectively, ie, for horses at 8, 9, 10 and 11 years old. At 10 years of age, the probability of still remaining was 43, 44, 45 and 50% for horses beginning at 4, 5, 6 and 7 years old, respectively, ie, after 7, 6, 5 and 4 years in competition. The half-lives (50% of horses still present in competition) decreased with age at first start from 6.1 years for horses starting at 4, to 3.5 for horses starting after 8 years (table II). The decrease was greatest between horses starting at 4 years old and those starting at 5 years old (0.8 year) and reduced to 0.1 year between 8 and 9 years old at first start.

The hazard function curves (fig 3) were increasing and the increase accelerated in the last years. This acceleration was in two steps: the first after 4 years in

Fig 2. Survival function of the length of jumping life: number of years in competition (x axis) versus probability (y axis).

Table II. Estimates of half-lives in function of age at first start.

Age at first start		half-life
4	years	6.1
5	years	5.3
6	years	4.5
7	years	4.0
8	years	3.6
9	years +	3.5
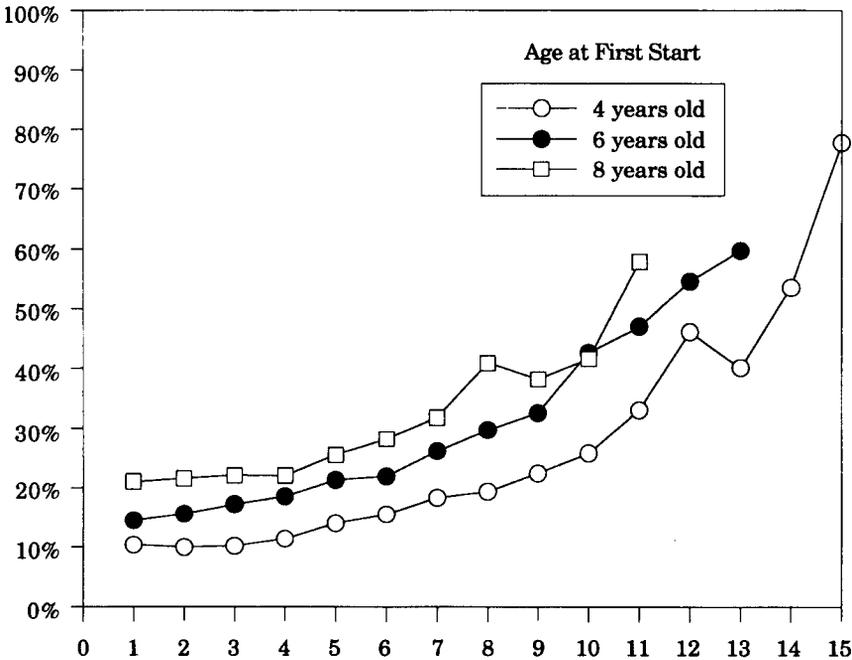
Fig 3. Hazard function of the length of jumping life: number of years in competition (x axis) versus probability (y axis).

competition and the second, more rapid one, after 9 years. The culling rate was smaller for horses that began earlier.

'Year' effect

The 'calendar year' effect was assumed to represent the variation in population size owing to herd management. Jumping is becoming more and more popular and the number of horses entering a show increased by 7% per year. The climatic variations and the evolution of management technology may also influence the length of competitive life. However, the censoring process explained the major part of the

variations and the preceding influences were hidden. Indeed, when the true date of culling of a horse were not known, he was considered as having failed when he did not appear between his last year of performance and the last year of recorded data. In case of a temporary interruption, the probability of appearing again decreased when the last year of performance approached the last year of data recording. This explains the higher relative culling rate for recent years (1.3 for 1989 and 1.6 for 1990). On the other hand, the first calendar years only included data from young horses, with a lower expected culling rate. This explains the low relative culling rate for 1972–1975 (0.6–0.8). Therefore, the year effects were likely to be more closely related to the structure of the data set than to environmental factors and were consequently difficult to interpret.
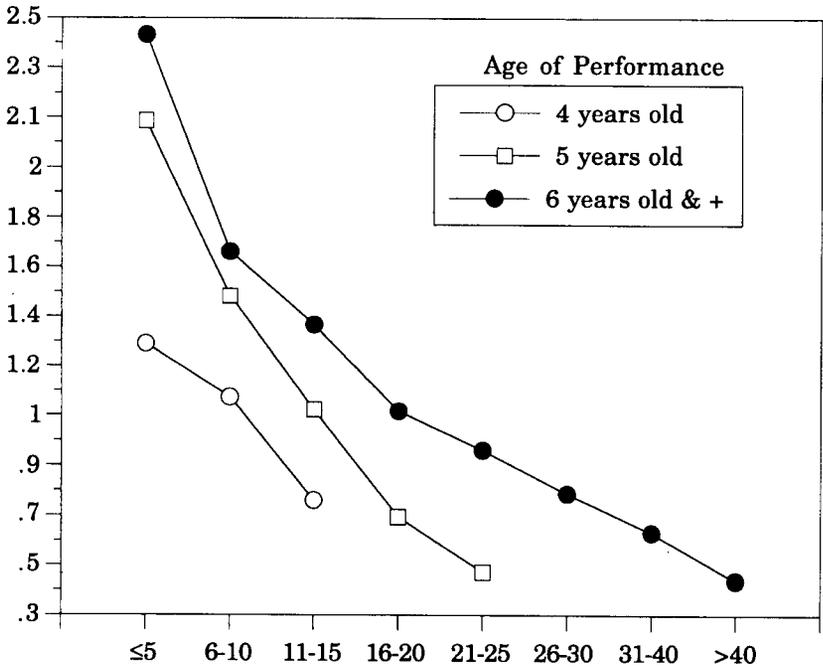
'Number of starts' effect

The relative culling rate associated with the 'number of starts' effect always decreased when the number of events increased (fig 4). This effect was only estimated with the auxilary model in order to obtain a correct adjustment for the level of performance. A high number of starts was probably not the reason for high longevity but rather an indication of good health and of the desire to continue jumping competitions. The 'number of starts' effect was moderate at 4 years old. It was more pronounced for horses aged 6 years and more. The effect was not linear: the decrease in the culling rate was more pronounced for a small numbers of starts.

'Level of performance' effect

After 6 years of age, the influence of the level of performance was clear: the better the horse, the greater his chance of continuing in competition (fig 5). The only exception was the slightly higher relative culling rate of horses with a performance rate higher than 130 but this difference was not significant. Horses that did not earn money had a strongly higher relative culling rate. A horse without earnings was 1.9 times more likely to be culled than an average horse with a performance rate between 90 and 100. This latter horse was 1.6 times more likely to be culled than a horse with a performance rate of 120–130. These results were expressed in terms of half-lives. For example, a horse that began the competition at 6 years old and had a performance level of 80–90 each year had a 1.5-year shorter half-life than a horse with a performance rate between 100 and 110 (5.4 versus 3.9 years). Owing to the large magnitude of performance effect, functional stayability is very different from true stayability. At 5 years old, the only significant difference concerned non-earning horses and good horses, with a smaller relative culling rate for the latter ones. The other horses had a similar relative culling rate. At 4 years old, the relative culling rate decreased as performance level increased but to a smaller extent than at 6 years old.

Breed effect

The relative culling rates of the three breeds of riding horses were very close: 0.90 for the Selle Français, 0.91 for the Anglo-Arab, 0.87 for the Cheval de Selle. The only

Fig 4. Number of annual starts effect: number of starts (x axis) versus relative culling rate (y axis).

significant difference was between Anglo-Arab and Cheval de Selle: an Anglo-Arab horse was 1.05 times more likely to be culled than a Cheval de Selle.

Thoroughbred and Trotteur Français typically start out as race horses and some of the unsuccessful ones later become jumpers: more than 50% of them began jumping at 6 years old. This new function was better tolerated by the Trotteur Français, whose relative culling rate was close to the Anglo-Arab (not significantly different), than by the Thoroughbred, which had a 1.26 times higher probability of being culled than the Trotteur Français. Two causes might explain this difference: either a prior racing career is less detrimental to a jumping career for trotters than for Thoroughbreds or trotters have a greater innate ability for tolerating the rigors of jumping competition.

Ponies and Arabs did not have jumping as a first objective and their high relative culling rate (1.2) might be the expression of their occasional use in competitions for horses.

Sire effect

The estimate of the γ parameter was 38.73. The expectation and variance of $w = \exp(s)$ were 1 and 0.0258, respectively, and the expectation and variance of s were -0.0130 and 0.0261. A phenotypic variance of the trait was needed to
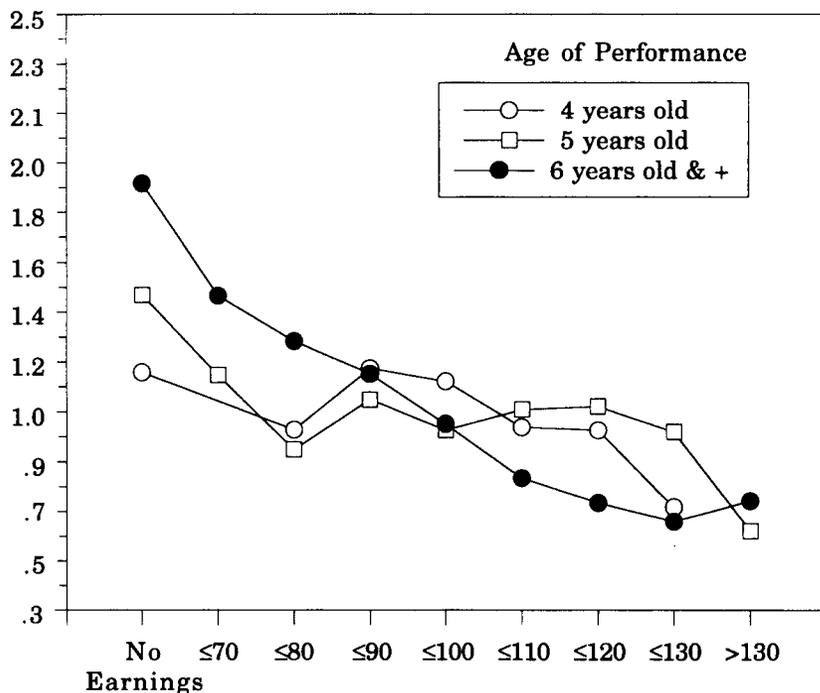
Fig 5. Level of jumping performance effect: performance rate based on Log(earnings), $\mu = 100, \sigma = 20$ (x axis) versus relative culling rate (y axis).

provide a corresponding heritability. This variance was difficult to define because the design of the explanatory variables was also dependent on time. In order to provide an estimate, taking into account age at record and age at first start effects, the variance of Log(t) varied from 0.5511 to 0.6023 according to the age at first start. The corresponding heritability was near 0.18. The mean of the distribution of the sire effects was -0.0273, and the standard deviation was 0.0485. The maximum was 0.2037 and the minimum was -0.3490. For example, the half-life difference between the progeny of the best and the worst sires was more than 2 years, if they started at 5 years old (respectively, 6.9 and 4.5 years). This difference was 0.4 year between offspring from a sire at $+1$ standard deviation and -1 standard deviation from the mean. The ratio of their hazards was 1.1. The genetic variability of the trait appeared to be particularly interesting. The heritability estimation was rather high compared to that obtained in dairy cattle (8.5%) by Ducrocq (1988b).

To provide an estimate of the genetic relationship between length of competitive life and jumping capacity, the correlation between breeding value estimates of sires for the two traits was computed. The sire breeding values for jumping capacity were obtained by an index based on the performances of the progeny. The correlation was -0.06, ie, close to zero or slightly favorable, between functional stayability, adjusted for level of performance, and jumping ability.

DISCUSSION AND CONCLUSION

This preliminary study identified some of the main factors influencing length of competitive life for jumping horses. The length of jumping life remains a trait difficult to define, because of the 'amateur' status of this sport on the one hand and because of the availability of data on the other.

An annual measure is in good agreement with the seasonal organization of competitions. However, is the criterion of a year's worth of performances really satisfied when the horse starts in only a few events? An alternative would be to require a minimum number of starts. Another possibility would be to define the time scale in terms of number of starts. To answer these questions accurately, genetic and phenotypic correlations have to be estimated between these different measures of the same trait with a multiple trait approach.

The data do not provide the exact date of the culling decision. The reason for the absence of horses from show jumping is not known, and is always considered as a true failure. This makes the interpretation of the 'year' effect unclear. In fact, the probability of being culled is dependent on the censoring probability. The closer the date of censoring, the higher is the probability for a horse to be considered as failed, because this horse does not have the opportunity to temporarily interrupt his jumping career. To minimize this problem, a better description of the censoring process is needed.

The characterization of the influence of jumping capacity also addresses several problems. It is not possible to clearly distinguish the respective proportions attributable to stayability and jumping ability in the relationship between annual earnings, number of starts in the year and length of active life within a year. The log(earnings) is indeed correlated with the number of starts, but also with the specific ability for jumping. This correlation is equal to 0.70 for horses aged 6 years and more. Moreover, this relation is not linear, but rather a logarithmic one. The number of starts is related to the length of life in the time interval considered (the mean number of starts for horses failing in a year is 7.3, against 15.6 for horses alive). And the jumping ability is also related to the number of annual starts: the better a horse is, the more he is used. The solution proposed here divides the influence of jumping ability on longevity between total earnings and the number of starts. Some other strategies are possible, based on earnings per start (correlation of 0.35 with the number of starts from 6 years old) possibly regressed on the number of starts, or based on different measures of sport capacity according to the number of starts. It remains critical to test the validity of each model. The likelihoods are always larger when the effect of the number of starts is included (the likelihood of the model with starts and earnings is better than with earnings alone) because the number of starts is a partial measure of time spent in the year and, consequently, of the existence of a culling. But the number of starts does not determine culling, it is only a consequence of culling. On the other hand, not adjusting for the level of performance would change the trait analyzed and increase its heritability because it would then approach the heritability of jumping ability, which is a major factor of length of competitive life. Finally, to confirm the genetic correlation between jumping ability and functional stayability, a multiple trait model is needed with a simultaneous estimation of the sire effects.

Nevertheless, the main results of this study are encouraging. The expected life of horses that began jumping early is the highest. The percentage of horses found at 9 or 10 years, the optimal age for performance, is almost constant, whatever their age of first start (4, 5 or 6 years). Good young horse management, with good rules for the competition of young horses that restricts the number of events, has no adverse effect on the length of their life, and produces horses with a better jumping capacity (Tavernier, 1992). According to the genetic correlation between early and mature performance (Tavernier, 1992), it is important to favor early selection of horses in competition on their early performances. A large majority (83%) of horses begins competition between 4 and 6 years of age: 40% at 4 years, 28% at 5 years and 15% at 6 years. The better stayability of horses that begin at 4 years of age is not only due to the benefit of their youth (the hazard function is increasing) but their relative culling rate becomes smaller, at the same age and until 13 years old, than that of horses that began at 5 years and especially at 6 years (differences after 13 years old are difficult to interpret because the standard deviations of estimates are large owing to the small size of the remaining population). Horses that began competition early have a true advantage that could be explained in two ways: either horses began at an early age because they showed good growth and health, or their learning of show jumping was better in the specific events for young horses, which then guarantees a long life. To reach his optimal capacity a horse has to learn the difficult sport of jumping, involving a long training period. He also needs to preserve his physical strength.

IMPLICATION

From a genetic improvement point of view, the length of jumping life is difficult to include in the selection objective: the heritability is low and the time needed to obtain enough information on the progeny of a sire is long (a fully informative observation is obtained when a horse has failed). Nevertheless, according to the genetic correlation obtained between length of competitive life and jumping capacity, selection on jumping is not expected to decrease the robustness of the horse.

Moreover, sires with poorer breeding values for the length of their jumping life may be detected. A medical and practical analysis of such a sire may reveal particular diseases and favor their genetic study. An evaluation of breeding value with an animal model, in addition to the present evaluation on performances (earnings), will also give important information for selecting stallions following their own jumping performance.

REFERENCES

Cox DR (1972) Regression models and life tables (with discussion). *J R Statist Soc* B 34, 187–280

Cox DR (1975) Partial likelihood. *Biometrika* 62, 269–276

Ducrocq V, Quaas RL, Pollak EJ, Casella G (1988a) Length of productive life of dairy cows. 1. Justification of a weibull model. *J Dairy Sci* 71, 3061–3070

Ducrocq V, Quaas RL, Pollak EJ, Casella G (1988b) Length of productive life of dairy cows. 2. Variance component estimation and sire evaluation. *J Dairy Sci* 71, 3071–3079

Kalbfleisch JD, Prentice RL (1980) *The Statistical Analysis of Failure Time Data*. Wiley, New York

Kaplan EL, Meier P 1958. Non parametric estimation from incomplete observations. *J Am Stat Assoc* 53, 457–469

NAG (1991) *NAG Fortran Library Guide, Mark 15*. NAG LTD, Oxford

Prentice RL, Gloeckler LA (1978) Regression analysis of grouped survival data with application to breast cancer data. *Biometrics* 34, 57–67

Tavernier A (1992) Is the performance at 4 years in jumping informative for later results? In: *EAAP Madrid, Horse Commission*, 13–17 September