

Boosting the First-Hitting-Time Regression Model

Vegard Stikbakke

December 18, 2018

Abstract

Empty.

Acknowledgements

Empty.

Contents

Abstract	i
Acknowledgements	iii
Contents	v
List of Figures	vii
List of Tables	ix
1 Introduction	1
2 Survival analysis	3
2.1 Survival data	3
2.2 Survival data likelihood regression setup	4
2.3 Proportional hazards regression	5
2.4 First hitting time models or threshold regression	6
3 Statistical boosting	11
3.1 AdaBoost	11
3.2 Boosting as Functional Gradient Descent	12
3.3 Gradient descent	13
3.4 Gradient boosting: Functional gradient descent	13
3.5 L_2 Boost	15
3.6 Component-wise gradient boosting	15
3.7 Multidimensional boosting: Component-wise boosting of a mul- tivariate loss function	17
3.8 The importance of stopping early	17
3.9 Selecting m_{stop}	19
4 First hitting time boost	23
4.1 Algorithm	23
Appendices	25
A Appendix 1: Differentiating the IG FHT	27
Bibliography	31

List of Figures

List of Tables

Chapter 1

Introduction

sec:intro

In this thesis, we work with boosting for regression in the first hitting time model. First hitting time is a model in survival analysis which serves as an alternative to the proportional hazards model, typically known as Cox regression. Developments in FHT regression are relatively recent, and there has to our knowledge been no attempt at tackling it in the high-dimensional case, in which boosting is an appropriate choice of method.

Chapter 2

Survival analysis

2.1 Survival data

Survival analysis is the field of studying lifetime and time-to-event data. An overview of modelling survival data is Aalen et al. (2008). We look at a stochastic variable $T > 0$ which is the time to event. To observe such data in real life, we must wait until the event actually happens. This might in some cases never happen, or it might take a very long time. One example is a clinical trial of n patients who have been treated for some disease, and where T_i , $i = 1, \dots, n$, is the time until they relapse. Typically these trials are for a set amount of time, say, until τ . Luckily, not every patient relapses during that time, and so their time of relapse T_i is not observed. We could throw away these observations, but we at least know that they survived until $t = \tau$. We therefore work with the concept of incomplete data, which we call *censored* data. An observed lifetime \hat{T} is censored if the actual lifetime T is larger than \hat{T} . We can say that we have a censoring mechanism which works such that the observed $\hat{T} = \min(T, C)$, where C is a censoring time. In the clinical trial example mentioned, $C = \tau$. We also need a censoring indicator, $D = I(\hat{T} = T)$, indicating if we have observed the actual event.

The survival function $S(t)$

In survival analysis, one of the things we are interested in is the survival function. The survival function $S(t)$ is the probability of surviving until time t ,

$$S(t) = \Pr(T > t) = 1 - \Pr(T \leq t) = 1 - F(t).$$

Here $F(t)$ is the familiar cumulative distribution function. If the derivative $f(t)$ of $F(t)$ exists, the lifetime T has probability distribution function $f(t)$.

The hazard function $\alpha(t)$

We are also interested in the hazard function. This is the probability of the event happening in a given small interval at time t , conditioned on the event not having happened yet. More formally, the hazard function is defined as a limit of this probability as the size of the interval goes to zero,

$$\alpha(t) = \lim_{\epsilon \rightarrow 0} \frac{\Pr(T < t + \epsilon | T > t)}{\epsilon}.$$

The hazard function $\alpha(t)$ is then the chance of the event happening at time t , if it has not happened yet. Estimation of the hazard function is hard, and we do not achieve the usual \sqrt{n} convergence.

Note that

$$\Pr(T < t + \epsilon | T > t) = \frac{\Pr(T < t + \epsilon, T > t)}{\Pr(T > t)} = \frac{F(t + \epsilon) - F(t)}{S(t)},$$

and inserting this into the hazard rate yields

$$\alpha(t) = \frac{1}{S(t)} \lim_{\epsilon \rightarrow 0} \frac{F(t + \epsilon) - F(t)}{\epsilon} = \frac{f(t)}{S(t)} = \frac{-S'(t)}{S(t)}, \quad (2.1) \quad \boxed{\text{\{eq:hfs\}}}$$

where the probability distribution function $f(t)$ is obtained by its limit definition, and we note that $S'(t)$ is the derivative of $1 - F(t)$, which is $-f(t)$. By integrating the hazard from 0 to time t , we get the cumulative hazard function $A(t) = \int_0^t \alpha(s) ds$,

$$A(t) = - \int_0^t \frac{S'(s)}{S(s)} ds = - \int_0^t \frac{\frac{dS}{df}}{S(s)} ds = - \int_0^t \frac{1}{S(s)} ds = - \log(S(t)). \quad (2.2) \quad \boxed{\text{\{eq:cumulative-hazard\}}}$$

Given censored survival data $(t_i, d_i), i = 1, \dots, n$, we introduce the at-risk function $Y(t)$, which is equal to the number of individuals still at risk at time t ,

$$Y(t) = \#\{t: t_i \geq t\},$$

where $\#(\cdot)$ is the counting operator over a set. We may then estimate the survival function $S(t)$ by the Kaplan-Meier estimator

$$\hat{S}(t) = \prod_{t_i \leq t} 1 - \frac{d_i}{Y(t_i)},$$

and the cumulative hazard function $A(t)$ by the Nelson-Aalen estimator

$$\hat{A}(t) = \sum_{t_i \leq t} \frac{d_i}{Y(t_i)}.$$

2.2 Survival data likelihood regression setup

Given survival data with covariates, (t_i, d_i, \mathbf{x}_i) , and parameterized functions $S(t|\mathbf{x}_i, \boldsymbol{\beta})$ and $f(t|\mathbf{x}_i, \boldsymbol{\beta})$ corresponding to a survival distribution, where $\boldsymbol{\beta} = (\beta_1, \dots, \beta_p)$ is a regression coefficient, we want to set up a likelihood for the data. We assume that the data is independent and identically distributed. If the event has occurred, the indicator d_i is 1. We can then use the information about the lifetime distribution, such that the single individual i contributes

$$f(t_i|\mathbf{x}_i) \quad (2.3) \quad \boxed{\text{\{eq:f\}}}$$

to the likelihood. If the event has not occurred, the observation is censored, and d_i is 0. In this case, we do not have the actual lifetime, and so we cannot use the lifetime distribution, but we must rather use the survival distribution. Therefore this observation contributes

$$S(t_i|\mathbf{x}_i) \quad (2.4) \quad \boxed{\text{\{eq:S\}}}$$

to the likelihood. Of course, since an observation can only be either censored or not censored at the same time, δ_i is either 0 or 1. If the event has occurred, d_i is 1, and then $1 - d_i$ is 0. Similarly, if the event has not occurred and the event is censored, then d_i is 0, and then $1 - d_i$ is 1. This allows us to take the product of (2.3) and (2.4) where we take these to the power of δ_i and $1 - \delta_i$, respectively, so that a single observation contributes

$$f(t_i|\mathbf{x}_i)^{\delta_i} S(t_i|\mathbf{x}_i)^{1-\delta_i}$$

to the likelihood. Since we assume the observations to be independent, the likelihood of the observed sample as a whole is the product of the single likelihoods. The complete likelihood becomes

$$L(\beta) = \prod_{i=1}^n f(t_i|\mathbf{x}_i, \beta)^{d_i} S(t_i|\mathbf{x}_i, \beta)^{1-d_i}. \quad (2.5) \quad \boxed{\{\text{eq:surv-lik}\}}$$

Since it is more convenient to work with the log likelihood, we derive this as well,

$$\begin{aligned} l(\beta) &= \log L(\beta) \\ &= \sum_{i=1}^n [d_i \log f(t_i|\mathbf{x}_i, \beta) + (1 - d_i) \log S(t_i|\mathbf{x}_i, \beta)]. \end{aligned} \quad (2.6) \quad \boxed{\{\text{eq:surv-lik}\}}$$

Note that since $\log S(t) = -A(t)$ (2.2) and $f(t) = \alpha(t)S(t)$ (2.1), this further simplifies to

$$l(\beta) = \sum_{i=1}^n [d_i \log \alpha(t_i|\mathbf{x}_i, \beta) - A(t_i|\mathbf{x}_i, \beta)].$$

2.3 Proportional hazards regression

So far we have not introduced covariates. How may we use a covariate vector \mathbf{x} in modelling, say, the hazard rate? A very common model to choose here is that of a proportional hazards model,

$$\alpha(t|\mathbf{x}) = \alpha_0(t)r(\mathbf{x}|\beta), \quad (2.7) \quad \boxed{\{\text{PH}\}}$$

where $\alpha_0(t)$ is an *unspecified* baseline hazard function shared between all individuals, and $r(\mathbf{x}|\beta)$ is a so-called relative risk function parameterized with regression coefficient $\beta = (\beta_1, \dots, \beta_p)$. We choose $r(\mathbf{x})$ such that it is appropriately normalized, meaning $r(\mathbf{0}) = 1$. A vital assumption here is that the covariates are fixed in time. With this setup, it turns out that we can do regression without specifying the baseline hazard. This is a major advantage, because we then do not have to think about modelling effects in time. Given data $(t_i, d_i), i = 1, \dots, n$, we may set up a so-called partial likelihood. For all observations $i = 1, \dots, n$ with $d_i = 1$, we know that there is an event at time t_i . The probability of the event happening for some individual j is the hazard, i.e., the instantaneous probability of that individual at that time, divided by the sum of all such hazards for those individuals still alive. Assuming that observations are independent and identically distributed, the partial likelihood

for the data is then the product of all such ratios,

$$\text{pl}(\beta) = \prod_{d_i=1} \frac{\Pr(\text{event happens to } i \text{ at time } t_i)}{\sum_{j \in R(t_i)} \Pr(\text{event happens to } j \text{ at time } t_i)} = \prod_{d_i=1} \frac{\alpha_0(t_i) r(\mathbf{x}_i)}{\sum_j \alpha_0(t_i) r(\mathbf{x}_j)},$$

where we see that the baseline hazard will cancel out, and we are left with just the relative risk functions.

The most common choice, by far, for $r(\mathbf{x})$ is the Cox model (Cox, 1992),

$$r(\mathbf{x}) = \exp(\mathbf{x}^T \beta).$$

This is an attractive model because the effect of a unit increase in an element of β has a nice interpretation. Assume we have two covariates \mathbf{x}_1 and \mathbf{x}_2 , and that \mathbf{x}_2 is equal to \mathbf{x}_1 except for in element j , where $x_{2j} = x_{1j} + 1$. Then the ratio of the two hazard rates becomes

$$\frac{\exp(\mathbf{x}_2^T \beta)}{\exp(\mathbf{x}_1^T \beta)} = \exp((\mathbf{x}_2 - \mathbf{x}_1)^T \beta) = \exp(\beta_j).$$

Cox regression is used a tremendous amount in applied research.

Cox regression example

Lorem ipsum.

The proportional hazards assumption

When we say (2.7), that $\alpha(t|\mathbf{x}) = \alpha_0(t)r(x|\beta)$, we make the proportional hazards (PH) assumption: We assume that the ratio between the hazard function of two individuals is the same *at all times*. This is a very large assumption to make, and in practice, it is very often not the case. One way to test this assumption for a covariate $j = 1, \dots, p$, is to fit a model $r(x_j) = \exp(f(x_j))$, where $f(\cdot)$ is some spline regression function, and plot it against x_j .

Robustness of Cox when the PH assumption is violated

Although the PH assumption is often not valid, in practice, Cox regression tends to work well.

CITATION NEEDED

2.4 First hitting time models or threshold regression

So far we have not thought much about how a time-to-event is generated. Instead, we have modelled the hazard rate directly. We have simply said that we have stochastic lifetimes. At one time, an individual is alive, and at a slightly later time, it is perhaps dead. One way to think about how these times are generated is to imagine that each individual has an underlying stochastic process, a health process $Y(t)$, say. Since the process is a function of time, it has a non-negative domain, $t \geq 0$. This health process is not observable, but when it hits a certain boundary set \mathcal{B} , the individual dies. \mathcal{B} is also called a barrier or a threshold, depending on what kind of set it is, and what association one wishes to invoke. The lifetime T , then, becomes the time it takes for the

health process $Y(t)$ to enter the boundary set \mathcal{B} . In general, the health process $Y(t)$ takes values in a set \mathcal{Y} , with an initial value $y_0 = Y(0)$. The barrier is a subset of this set of values, $\mathcal{B} \in \mathcal{Y}$, with the initial health process value $y_0 \notin \mathcal{B}$. In other words, the lifetime is

$$T = \operatorname{argmin}_t Y(t) \in \mathcal{B}.$$

First hitting time (FHT) models were introduced in Whitmore (1986), and a good reference paper on the topic is Lee and Whitmore (2006). Note that these authors use the term threshold regression. We have, together with Caroni (2017), chosen to not use this term, as it is also the name of an already established, and quite different, field of econometrics. FHT models have been applied to many different fields, including medicine, engineering, and economics. They may describe the survival time of a transplant patient, the duration time of a strike, the failure time of an engineering system, and so on.

The first hitting time model framework is highly flexible. We have flexibility both in choice of process, boundary and initial value. The most important part is the stochastic process. Examples include Wiener processes, Markov chains, Bernoulli processes, and Gamma processes. We choose to use the Wiener process, because it turns out that it yields a fully parametric regression model.

Wiener process

Let $W(t)$ be a continuous stochastic process defined for $t \in [0, \infty)$, taking values in \mathbb{R} , and with initial value $W(0) = 0$. If W has increments that are independent and normally distributed with

$$E[W(s+t) - W(t)] = 0 \text{ and } \operatorname{Var}[W(s+t) - W(t)] = s,$$

we call W a Wiener process. Its position at time t is always a Gaussian distribution $N(0, t)$ (Aalen et al., 2008). To increase the flexibility of the Wiener process, we can let

$$Y(t) = y_0 - \mu t + \sigma W(t), \tag{2.8}$$

{wiener}

which is called a Wiener process with initial value y_0 , drift coefficient μ , and diffusion coefficient σ . Introductions to many aspects of Wiener processes are found in Cox and Miller (1965). Figure shows simulations of some Wiener process paths.

there should be a figure here!

FHT with Wiener process leads to Inverse Gaussian

If we choose the stochastic process to be a Wiener process like in (2.8), and we let the boundary be the non-positive numbers, $\mathcal{B} = (-\infty, 0]$, then the lifetime is the time it takes for the process to first reach a non-positive value,

$$T = \operatorname{argmin}_t Y(t) \leq 0. \tag{2.9}$$

Note that since the Wiener process is continuous, there will not be a difference between \leq and $<$.

This is a very conceptually appealing model, because it assumes that individuals might have different initial levels, and that also the drift might be different between individuals. It is also attractive because it has closed-form probability and cumulative density functions, and its likelihood is computationally simple. There are no restrictions on the movements of the process, meaning, it is non-monotonic. If we do want a monotonic restriction on the movement of the process, we may use a gamma process (Lee and Whitmore, 2006).

It can be shown that the first hitting time of the Wiener process follows an inverse Gaussian distribution (Chhikara, 1988), with probability distribution function

$$f(t|y_0, \mu, \sigma^2) = \frac{y_0}{\sqrt{2\pi\sigma^2 t^3}} \exp \left[-\frac{(y_0 + \mu t)^2}{2\sigma^2 t} \right], \quad (2.10)$$

{eq:ig-pdf}

and cumulative distribution function

$$F(t|\mu, \sigma^2, y_0) = \Phi \left[-\frac{(\mu t + y_0)}{\sqrt{\sigma^2 t}} \right] + \exp \left(-\frac{2y_0\mu}{\sigma^2} \right) \Phi \left[\frac{\mu t - y_0}{\sqrt{\sigma^2 t}} \right]. \quad (2.11)$$

{eq:ig-cdf}

TO DO!

See Appendix for the mathematical derivation. Note that if the drift μ is positive, then it is not certain that the process will ever reach 0. Hence the probability distribution function in (2.10) is improper. In this case, the probability of the time not being finite is (Cox and Miller, 1965)

$$\Pr(T = \infty) = 1 - \Pr(T < \infty) = 1 - \exp(-2y_0\mu).$$

Since we in survival analysis prefer working with the survival function $S(t) = 1 - F(t)$ rather than the cdf $F(t)$, we note that it becomes

$$S(t|\mu, \sigma^2, y_0) = \Phi \left[\frac{\mu t + y_0}{\sqrt{\sigma^2 t}} \right] - \exp \left(-\frac{2y_0\mu}{\sigma^2} \right) \Phi \left[\frac{\mu t - y_0}{\sqrt{\sigma^2 t}} \right], \quad (2.12)$$

{eq:ig-surv}

where $\Phi(x)$ is the cumulative distribution function of the standard normal, i.e.,

$$\Phi(x) = \int_{-\infty}^x \exp \left(-\frac{y^2}{2} \right) / \sqrt{2\pi} \, dy, \quad (2.13)$$

and in (2.12) we used the fact that $1 - \Phi(-x) = \Phi(x)$, since the standard normal distribution is symmetric around 0.

The inverse gaussian is overdetermined if the health process is latent

There are three parameters in the inverse Gaussian distribution, namely y_0 , μ and σ . We observe, however, that both the pdf $f(t|y_0, \mu, \sigma^2)$ in (2.10) and the survival function $S(t|\mu, \sigma^2, y_0)$ in (2.12) only depend on these parameters through μ/σ and y_0/σ . Hence, there are only two free parameters. In other words, we can without loss of generality fix one parameter, for instance set σ equal to 1. This is the conventional way to do it (Lee and Whitmore, 2006).

The shape of the hazard rate

The hazard rate is obtained from $\alpha(t) = f(t)/S(t)$ (2.1).

Regardless of initial value, this converges to the same limiting hazard. If y_0 is close to zero, we essentially get a decreasing hazard rate. If y_0 is far from zero, this gives an essentially increasing hazard rate. If y_0 is somewhat inbetween, we get a hazard rate which first increases and then decreases (Aalen et al., 2008).

Comparison of hazard rates

Of particular interest might be looking at the ratio between two hazard rates, that is, one hazard divided by the other. We might for example look at it when the drift μ is the same, but the initial level y_0 is different. Then the hazard ratio is strongly decreasing. It is also of interest to do the converse, that is, look at the hazard ratio when the initial level is the same, but the drift is different. The result here is quite different. The ratio of the hazards has a “bathtub” shape, which levels off at a later time (Aalen et al., 2008). Keep in mind here that levelling off means getting to proportional hazards.

The hazard function converges to

$$\lim_{t \rightarrow \infty} \alpha(t) = \frac{1}{2} \left(\frac{\mu}{\sigma} \right)^2 = 0.5\mu^2 \quad (2.14)$$

We see that the FHT framework with a Wiener process is a highly flexible parametric model for survival analysis. Indeed, much more flexible than Cox regression, since the hazard ratios in Cox are all confined to be constant over time.

Regression

We may introduce effects from covariates by allowing μ and y_0 to depend on covariates \mathbf{x} and \mathbf{z} . A simple and much used model is to simply use the identity link function for the drift μ , and to use the logarithm link function for the initial level y_0 , since it must be positive.

$$\begin{aligned} \mu &= \boldsymbol{\beta}^T \mathbf{x} \\ \ln y_0 &= \boldsymbol{\gamma}^T \mathbf{z} \end{aligned} \quad (2.15)$$

{eq:coeffs}

Here $\boldsymbol{\beta}$ and $\boldsymbol{\gamma}$ are vectors of regression coefficients. Note that we may let \mathbf{x} and \mathbf{z} share none, some, or all elements. We will discuss consequences of this later.

Inserting the pdf (2.10) and the survival function (2.12) into the loglikelihood (2.6), we get that the loglikelihood of a survival data set with the inverse gaussian FHT model is

$$\begin{aligned} l(y_0, \mu, \sigma) &= \sum_{i=1}^n \delta_i \left(\ln y_0 - \frac{1}{2} \ln(2\pi\sigma^2 t_i^3) - \frac{(y_0 + \mu t_i)^2}{2\sigma^2 t_i} \right) \\ &\quad + (1 - \delta_i) \ln \left(\Phi \left(\frac{\mu t_i + y_0}{\sqrt{\sigma^2 t_i}} \right) - \exp \left(-\frac{2y_0\mu}{\sigma^2} \right) \Phi \left(\frac{\mu t_i - y_0}{\sqrt{\sigma^2 t_i}} \right) \right). \end{aligned} \quad (2.16)$$

{eq:loglik}

Fitting an IG FHT model

At the moment, the standard for fitting an inverse gaussian FHT model to survival data is to use numerical likelihood maximization Caroni (2017). A few software packages exist for doing this, and one of these for R R Core Team (2013) is the **threg** package Xiao et al. (2015). There does not exist any method to fit a *regularized* model at the moment. **This is the main focus of my thesis.**

Example of application

Lorem ipsum some example. Just use numerical maximization.

Identification problems

Chapter 3

Statistical boosting

Boosting is one of the most promising methodological approaches for data analysis developed in the last two decades (Mayr et al., 2014). It is now a staple part of the statistical learning toolbox because it is a flexible tool for estimating interpretable statistical models. Boosting, however, originated as a black box algorithm in the fields of computational learning theory and machine learning, not in statistics.

Computer scientists Michael Kearns and Leslie Valiant, who were working on computational learning theory, posed the following question: Could any weak learner be transformed to become a strong learner? (Kearns and Valiant, 1989) A weak learner, sometimes also simple or base learner, means one which has a low signal-to-noise ratio, and which in general performs poorly. For classification purposes it is easy to give a good example: A weak learner is one which performs only slightly better than random uniform chance. In the binary classification setting, then, it would only perform slightly better than a coin flip. For regression, a weak learner is for example a linear least squares model of only one variable, and having only a small parameter effect for that variable. Meanwhile, a strong learner should be able to perform in a near-perfect fashion, for example attaining 99% accuracy on a prediction task. I will first attend to give a summary of the history of boosting, starting with AdaBoost (Freund and Schapire, 1996), which proved that the answer to the original question above was yes. For another overview, consult also the literature review article by Mayr et al. (2014).

3.1 AdaBoost

The original AdaBoost, also called Discrete AdaBoost (Freund and Schapire, 1996) is an iterative algorithm for constructing a binary classifier $F(\cdot)$. It was the first *adaptive* boosting algorithm, as it automatically adjusted its parameters to the data based on its performance. In the binary classification problem, we are given a set of observations $(\mathbf{x}_i, y_i)_{i=1, \dots, n}$, where $x \in \mathbb{R}^p$ and $y \in \{-1, 1\}$, i.e., positive or negative; yes or no. We want to find a rule which best separates these observations into the correct buckets, as well as being able to classify new, unseen observations \mathbf{x}_{new} of the same form. Some observations are hard to classify, whereas some are not. One way to look at binary classification is to imagine the p -dimensional space of the observations \mathbf{x} , and think of the classifier

as finding the line which best splits the observations into their corresponding label. Some observations are not at all close to the boundary, and so they are easily classified. Other observations, however, are close to the boundary. Freund and Schapire (1996) realized that one could assign a weight to each observation. First, assign equal weight to each observation. Then, use a weak learner $h(\cdot)$ to make an initial classifier, minimizing the weighted sum of misclassified points, which initially is a plain sum. After this initial classification, some points will be correctly classified, and some will be misclassified. Change the weights in a certain way such that the misclassified ones get increased weights, and normalize the weights afterwards. Based on the misclassification rate, calculate a weight α to give to this classifier. This results in the correctly classified observations having smaller weight than before. Currently, the classifier is $F_1(\cdot) = \alpha_1 h_1(\cdot)$. In the next iteration, make a new weak learner which minimizes the weighted sum of the observations and reweight observations accordingly as before. Again calculate a weight to give to this new classifier, and add it to the previous classifier, such that $F_2(\cdot) = \alpha_1 h_1(\cdot) + \alpha_2 h_2(\cdot)$. Continue iterating in this fashion until an iteration m . The resulting final classifier, the AdaBoost classifier, becomes $\hat{F}(\cdot) = F_m(\cdot) = \sum_{i=1}^m \alpha_i h_i(\cdot)$. It is a linear combination of the weak classifiers, and in essence a weighted majority vote of weak learners given the observations.

The AdaBoost algorithm often carries out highly accurate prediction. In practice, it is often used with stumps: Decision trees with one split. For example, Bauer and Kohavi (1999) report an average 27% relative improvement in the misclassification error for AdaBoost compared to a single decision tree. They conclude that boosting not only reduces the variance in the prediction error from using different training data sets, but that it also is able to reduce the average difference between the predicted and the true class, i.e., the bias. Breiman (1998) supports this analysis. Because of its plug-and-play nature and the fact that it never seemed to overfit (overfitting occurs when the learned classifier degrades in test error because of being too specialized on its training set), Breiman remarked that “boosting is the best off-the-shelf classifier in the world” (Hastie et al., 2009).

Something about overfitting?

In its original formulation, the AdaBoost classifier does not have interpretable coefficients, and as such it is a so-called black-box algorithm. This means that we are unable to infer anything about the effect of different covariates. In statistics, however, we are interested in models which are interpretable.

See some figure for a schematic overview of the algorithm.

3.2 Boosting as Functional Gradient Descent

While originally developed for binary classification, boosting is now used to estimate the unknown quantities in more general statistical models and settings. We therefore extend our discussion to a more general regression scheme. Let $D = \{x^{(i)}, y^{(i)}\}_{i=1, \dots, n}$ be a learning data set sampled iid from a distribution over the joint space $\mathcal{X} \times \mathcal{Y}$, where the input space is a possibly high-dimensional $\mathcal{X} \in \mathbb{R}^p$ and the output space is a low-dimensional space \mathcal{Y} . For the majority of applications, the output space \mathcal{Y} is one-dimensional and continuous, e.g., in the standard regression setting. (In the censored survival data setting, however, it is

two-dimensional, since we have $y^{(i)} = (t_i, d_i)$.) We are interested in estimating the expectation of Y given a realization x of X ,

$$\mathbb{E}[Y|\mathbf{X} = \mathbf{x}] = F(\mathbf{x}),$$

in an interpretable way such that we can quantify the relation between one or more predictor variables and the expectation of the response. We choose a model $g(\cdot)$ to model $F(\cdot)$, and assume that we can approximate $F(\cdot)$ sufficiently by an additive model

$$\hat{F}(\mathbf{x}) = \eta(\mathbf{x}) \quad (3.1)$$

where $F(\cdot)$ is some function and where $\eta: \mathcal{X} \rightarrow \mathbb{R}$ is an additive predictor

$$\eta(\mathbf{x}) = \beta_0 + \sum_{j=1}^J f_j(\mathbf{x}|\beta_j), \quad (3.2)$$

{eq:eta}

with a **constant intercept coefficient** β_0 , and additive effects $f_j(x|\beta_j)$.

To calculate the difference between the estimated outcome and the actual outcome, we need a loss function ρ . It measures the difference, or distance, between the true outcome $y^{(i)}$ and the additive predictor $\eta(x^{(i)})$. A loss function must be symmetric and convex. Examples of ρ are the absolute loss $|y - \eta(x)|$, which leads to a regression model for the median, and the quadratic loss (the L_2 loss), which leads to the usual regression model for the mean. Very often, the loss is derived from the **negative** log likelihood of the distribution of \mathcal{Y} , depending on the desired model. Keep in mind that the negative is used, due to the aim of *minimizing* the loss function. In the survival data setting, typical loss functions are ROC and Briar score (Bøvelstad and Borgan, 2011).

In this setting, Friedman et al. (2000) developed an algorithm for fitting an additive model, which he showed performs gradient descent in functional space. He further showed that AdaBoost performs this algorithm, for a particular exponential loss function. This provided a way of viewing boosting through a statistical lens, and connected the successful machine learning approach to the world of statistical modelling.

3.3 Gradient descent

Gradient descent is an optimization algorithm for minimizing a differentiable multivariate function F . The motivation behind the gradient descent algorithm is that in a small interval around a point \mathbf{x} , F is decreasing in the direction of the negative gradient at \mathbf{x} . Therefore, by moving slightly in that direction, F will decrease. Indeed, with a sufficiently small step length, gradient descent will always converge, albeit to a local minimum. For a schematic overview of the algorithm, see Algorithm 1.

3.4 Gradient boosting: Functional gradient descent

Consider the problem of finding an $F(\cdot)$ to minimize the empirical risk (3.9). Often we choose a parameterized model $F(\mathbf{X}; \beta)$. Finding the optimal β analytically might be infeasible. Enter the gradient descent algorithm. In this case, we look at a fixed input data set \mathbf{X} and let $F(\cdot)$ only be a function of

algo:grad-desc

Algorithm 1 Gradient descent

1. Start with an initial guess \mathbf{x}_0 , e.g. $\mathbf{x}_0 = \mathbf{0}$. Let $m = 1$.
2. Calculate the direction to step in, $\mathbf{g}_{m-1} = -\nabla F(\mathbf{x}_{m-1})$.
3. Let $\mathbf{h}_m = \nu \mathbf{g}_{m-1}$, where ν is the best step length, found by

$$\nu = \underset{\nu}{\operatorname{argmin}} \mathbf{x}_{m-1} + \nu \mathbf{g}_{m-1}$$

4. Let $\mathbf{x}_m = \mathbf{x}_{m-1} + \mathbf{h}_{m-1}$
5. Increase m , and go to step 2. Repeat until $m = M$.
6. The resulting final guess is $\mathbf{x}_M = \mathbf{x}_0 + \sum_{m=1}^M \mathbf{h}_m(\mathbf{x}_m)$

β . Thus we can use gradient descent to find an optimal β , in other words doing gradient descent in parameter space (Friedman, 2001). Boosting can be viewed as an optimization procedure in functional space. We first describe a naive way of doing this. Consider the function value at each \mathbf{x} directly as a parameter, and use gradient descent directly on these parameters. However, this does not generalize to unobserved values \mathbf{X} , and we are after all interested in the population minimizer of (??). We can instead assume a parameterized form for F , e.g.,

$$F(\mathbf{X}; \{\beta\}_{m=1}^M) = \sum_{m=1}^M \nu f(\mathbf{X}; \beta_m), \quad (3.3)$$

{eq:gradboost}

where $f(\mathbf{X}; \beta)$ is an additive model, but typically a simple one, i.e., a base learner as discussed previously. We would like to minimize a data based estimate of the loss, i.e. the empirical risk (3.9), and so would choose $\{\beta_m\}$ as the minimizers of

$$\underset{\{\beta_m\}_{m=1}^M}{\operatorname{argmin}} \overline{\operatorname{err}}(f(\mathbf{x}; \{\beta_m\})).$$

However, estimating these simultaneously may be infeasible. We can then use a greedy stagewise approach, where we at each step m choose the β_m which gives the best improvement, while not changing any of the previous $\{\beta\}_{k=1}^{m-1}$. This means that we look at the *generalized residuals* given the previous parameters. At each step m the current solution is

$$F_m = F_{m-1} + \nu f(\mathbf{x}; \beta_m),$$

where the parameters β_m are those in f minimizing the empirical risk when added to the fixed part F_{m-1} :

$$\beta_m = \underset{\beta}{\operatorname{argmin}} \overline{\operatorname{err}}(f(\mathbf{x}; \beta_k) + f(\mathbf{x}; \beta)).$$

The final model is then the sum of these terms, like in (3.3). To find β_m in each step here, we use gradient descent. This concludes the outline of a generic functional gradient descent algorithm. For a schematic overview, see Algorithm 2. Note that while we call this functional gradient descent (FGD), this is exactly the gradient boosting algorithm.

algo:fgd

Algorithm 2 Functional gradient descent

1. Initialize $F_0(\mathbf{x})$, e.g., by setting it to zero for all components. Select a base learner f .
2. Compute the negative gradient vector,

$$U_i = -\frac{\partial L(y_i, F_{m-1}(\cdot))}{\partial F_{m-1}(\cdot)}, \quad i = 1, \dots, N.$$

3. Estimate \hat{h}_m by fitting (\mathbf{X}_i, U_i) using the base learner h (like in the previous algorithm):

$$\beta_m = \operatorname{argmin}_{\beta} \sum_{i=1}^N L(u_i, h(\mathbf{x}_i; \beta))$$

This estimation can be viewed as an approximation of the negative gradient vector.

4. Let $f_m = \nu h(\cdot; \beta_m)$ and update $F_m(\cdot) = F_{m-1}(\cdot) + f_m$.
5. Repeat steps from 2 until $m = M$.

3.5 L_2 Boost

Friedman (2001) continued to investigate the topic, providing further insight into boosting. Based on his results, Bühlmann and Yu (2003) developed the L_2 Boost algorithm. It is a special case of the generic functional gradient descent (FGD) algorithm, where we choose the squared error loss to be the loss function,

$$L(y, F(\mathbf{x})) = \frac{1}{2} (y - F(\mathbf{x}))^2.$$

The negative gradient vector of the loss then becomes the residual vector,

$$\frac{\partial L(y, F(\mathbf{x}))}{\partial x_i} = (y - F(x_i)), \quad i = 1, \dots, n,$$

and hence the boosting steps become repeated refitting of residuals (Friedman, 2001; Bühlmann and Yu, 2003). With $\nu = 1$ and $M = 2$, this had been proposed already by (Tukey, 1977), who called it “twicing”. See Algorithm 3 for an overview.

They also prove some nice important theoretical results for L_2 Boost.

more on L_2 Boost!!

3.6 Component-wise gradient boosting

In high-dimensional settings, it might often be infeasible, if not impossible, and at least impractical, to use a base learner h which incorporates all p dimensions. Indeed, using least squares base learners, it is impossible, since the matrix which

algo:l2

Algorithm 3 L2Boost

1. Initialize $F_0(\mathbf{x})$, e.g., by setting it to zero for all components. Select a base learner h , such as ordinary least squares, stumps, or smoothing splines.
2. Compute the residuals

$$U_i = (y_i, F_{m-1}(x_i)), \quad i = 1, \dots, n$$

3. Estimate \hat{h}_m by fitting $(\mathbf{X}_i, U_i)_{i=1}^N$ using the base learner h :

$$\beta_m = \underset{\beta}{\operatorname{argmin}} \sum_{i=1}^N L(u_i, H(\mathbf{x}_i; \beta))$$

Note that $\hat{H}(\cdot; \beta_m)$ is then an estimate of the negative gradient vector.

4. Update $F_m(\cdot) = F_{m-1}(\cdot) + \nu H(\cdot; \beta_m)$.
 5. Repeat steps from 2 until $m = M$.
-

must be inverted is singular when $p > N$. Component-wise gradient boosting is a technique/algorithm which does work in these settings. First developed by Bühlmann and Yu (2003), and it has further been refined and explored, see e.g. Bühlmann (2006). The idea of the algorithm is to select a set of base learners, the most important property of which being that they are univariate: Each base learner is only a function of one component x_j of the data \mathbf{X} , i.e.,

$$h_j(\mathbf{x}) = h_j(x_j).$$

In each iteration, we fit the learners separately, and choose only the one which gives the best improvement to be added in the final model. The resulting model $F_m(\cdot)$ is then a sum of componentwise effects,

$$F_m(\mathbf{X}) = \sum_{j=1}^p f_j(x_j),$$

where

$$f_j(x_j) = \sum_{m=1}^M \mathbb{1}_{mj} h_j(x_j; \beta_m),$$

where $\mathbb{1}_{mj}$ is an indicator function which is 1 if component j was selected at iteration m and 0 if not. Hence this model is a GAM (??). Crucially, if we stop sufficiently early, we will typically perform variable selection. It is likely that some base learners have never been added to the final model, and as such those components in \mathbf{X} are not added. For a schematic overview of the algorithm, see Algorithm 4. In fact, Bühlmann believes that it is mainly in the case of high-dimensional predictors that boosting has a substantial advantage over classical approaches (Bühlmann, 2006).

algo:component-
gradboost

Algorithm 4 Component-wise gradient boosting

1. Start with an initial guess, e.g. $F_0 = \mathbf{0}$.
Specify a set of base learners $h_1(\cdot), \dots, h_p(\cdot)$.
2. Compute the negative gradient vector \mathbf{U} .
3. Fit $(\mathbf{X}_i, U_i)_{i=1}^N$ separately to every base learner to get $\hat{h}_1(x_1), \dots, \hat{h}_p(x_p)$.
4. Select the component k which best fits the negative gradient vector.

$$k = \operatorname{argmin}_{j \in [1, p]} \sum_{i=1}^N (u_i - \hat{h}_j(\mathbf{x}_i))^2$$

5. Update $F_m(\cdot) = F_{m-1}(\cdot) + \nu h_k(x_k)$
-

3.7 Multidimensional boosting: Component-wise boosting of a multivariate loss function

The above methods consider a loss function depending on one parameter: $L(\beta)$. In the boosting steps, one uses a gradient descent step with the loss function differentiated with respect to this one parameter. There are however many models which use loss functions of several variables. Schmid et al. (2010) extend the gradient boosting algorithm (Friedman, 2001) to such a setting. In this method, we take the partial derivative of the multivariate loss function with respect to each variable. At this point, there are several options for which path to take. We provide an overview of several of these.

Schmid et al. (2010) propose to, in each boosting iteration, do a boosting step in each dimension at a time, and to repeat this for all dimensions. They call this *cycling* through the parameters, hence the algorithm name. This also involves finding a vector $\mathbf{m}_{\text{stop}} = (m_{\text{stop},j})_{j=1,\dots,k}$ for each dimension $j \in 1, \dots, p$. The algorithm proposed in Schmid et al. (2010) does an additional cycle through of nuisance parameters after having cycled through all dimensions, in each boosting step. Mayr et al. (2012a) applies this to the GAMLSS model originally developed by Rigby and Stasinopoulos (2005), and this is the algorithm we reproduce. It is basically the same without nuisance parameters. The algorithm is as follows. Thomas et al. (2018) point out a problem with this and that is that the different $m_{\text{stop},j}$ parameters are not independent of each other, and hence have to be jointly optimized. The usually applied grid search for such parameters scales exponentially with the number of parameters k , and this can quickly become computationally very demanding.

3.8 The importance of stopping early

The number of iterations in the boosting procedure, M , is a tuning parameter. It acts as a regularizer.

algo:multidim-
boost

Algorithm 5 Cyclical multidimensional component-wise gradient boosting

1. Initialize the n -dimensional vectors $f_1^{[0]}, \dots, f_K^{[0]}$, with offset values, e.g. with $f_1^{[0]} = \mathbf{0}, \dots, f_K^{[0]} = \mathbf{0}$. Alternatively, one can use the maximum likelihood estimates as offset values.
 2. For each component $k = 1, \dots, K$, specify a base learner to use for each of the p components. The base learner takes one input variable and has one output variable. Examples include.
 3. Set $m = 0$.
 4. Increase m by 1.
 - a) Set $k = 0$.
 - b) Increase k by 1. If $m > m_{\text{stop},k}$, proceed to step 4 f). If not, compute the negative partial derivative $-\frac{\partial \rho}{\partial f_k}$ and evaluate at $\hat{f}^{[m-1]}(X_i) = (\hat{f}_1^{[m-1]}(X_i), \dots, \hat{f}_K^{[m-1]}(X_i))$, $i = 1, \dots, n$. This yields the negative gradient vector $U_k^{[m-1]} = (U_{i,k}^{[m-1]})_{i=1, \dots, n} := \left(-\frac{\partial}{\partial f_k} \rho(Y_i, \hat{f}^{[m-1]}(X_i)) \right)_{i=1, \dots, n}$.
 - c) Fit the negative gradient vector $U_k^{[m-1]}$ to each of the p components of \mathbf{X} separately (i.e. to each predictor variable) using the base learners specified in step 2. This yields p vectors of predicted values, where each vector is an estimate of the negative gradient vector $U_k^{[m-1]}$.
 - d) Select the component of \mathbf{X} which best fits $U_k[m-1]$ best according to a pre-specified goodness-of-fit criterion. For continuous variables, the R^2 measure should (?) be used. Set $\hat{U}_k^{[m-1]}$ equal to the fitted values of the corresponding best model fitted in the previous step.
 - e) Update $\hat{f}_k[m-1] \leftarrow \hat{f}_k^{[m-1]} + \nu \hat{U}_k^{[m-1]}$, where ν is a real-valued step-length factor.
 - f) Update nuisance parameter(s).
 - g) For $k = 2, \dots, K$, repeat steps. Finally, update $\hat{f}^{[m]} \leftarrow \hat{f}^{[m-1]}$.
 5. Iterate step 4 until $m > \max(m_{\text{stop},k})$ for all $k \in \{1, \dots, K\}$.
-

3.9 Selecting m_{stop}

The crucial tuning parameter in boosting is the number of iterations, m_{stop} . Stopping early enough performs variable selection and shrinks the parameter estimates toward zero. Left on its own, the parameters in boosting will converge towards the maximum likelihood parameters, i.e., maximizing the in-sample error. We are, on the other hand, after all interested in minimizing out-of-sample prediction error (PE). The prediction error for a given data set is a function of the boosting iteration m . What we want is therefore a good method for approximating $\text{PE}(m)$. This can be done in a number of ways. Many authors state that the algorithm should be stopped early, but do not go further into the details here. Common model selection criteria such as the Akaike Information Criteria (AIC) may be used, however the AIC is dependant on estimates of the model's degrees of freedom. Methods by Chang et al. (2010) try this. This is problematic for several reasons. For L_2 Boost, Bühlmann and Hothorn (2007) suggest that $\text{df}(m) = \text{trace}(B_m)$ is a good approximation. Here B_m is the hat matrix resulting from the boosting algorithm. This was, however, shown by Hastie (2007) to always underestimate the actual degrees of freedom. Mayr et al. (2012b) propose a sequential stopping rule using subsampling etc. We argue instead that cross-validation, a very common method for selection of tuning parameters in statistics, is a good method to use. It is flexible and easy to implement. It is somewhat computationally demanding, requiring several full runs of the boosting algorithm.

SOURCE?

K-fold cross-validation

K-fold cross-validation (Lachenbruch and Mickey, 1968), or simply cross-validation, is a general method commonly used for selection of penalty or tuning parameters. We will use it to approximate the prediction error. In cross-validation, the data is split randomly into K roughly equally sized folds. For a given fold k , all folds except k act as the training data in estimating the model. We often say that the k 'th fold is left out. The resulting model is then evaluated on the unseen data, namely fold k . This procedure is repeated for all $k = 1, \dots, K$. An estimate for the prediction error is obtained by summing over the test errors from evaluating the left-out fold. Let $\kappa(k)$ be the set of indices for fold k . The cross-validated estimate for a given m then becomes

$$\text{CV}(m) = \sum_{k=1}^K \sum_{i \in \kappa(k)} L((t_i, \delta_i), \theta_m; \mathbf{x}_i). \quad (3.4)$$

For each m , we calculate the estimate of the cross-validated prediction error $\text{CV}(m)$. We choose m_{stop} to be the minimizer of this error,

$$m_{\text{stop}} = \underset{m}{\text{argmin}} \text{CV}(m). \quad (3.5)$$

Typical values for K are 5 or 10, but in theory one can choose any number. The extreme case is $K = N$, called leave one out cross-validation, where all but one observation is used for training and one evaluates the model on the observation that was left out. In this case, the outcome is deterministic, since there is no randomness when dividing into folds.

Stratified cross-validation

When dividing an already small number of survival data observations into K folds, we might risk getting folds without any observed deaths, or in any case, very few. In stratified cross validation, we do not divide the folds entirely at random, but rather, try to divide the data such that there is an equal amount of censored data in each fold. As before, let $\kappa(k)$ be the set of indices for fold k . Divide the observed data into K folds, as with usual cross validation, to get an index set $\kappa_{\delta=1}(k)$ for a given k . Similarly, divide the censored data into K folds, obtaining $\kappa_{\delta=0}(k)$. Finally, $\kappa(k)$ is the union of these sets: $\kappa(k) = \kappa_{\delta=1}(k) \cup \kappa_{\delta=0}(k)$. For “real-life data sets like ours”, Kohavi (1995) illustrate that 10-fold stratified cross validation performs best.

Repeated cross-validation

The randomness inherent in the cross-validation splits has an effect on the resulting m_{stop} . This is true for boosting in general, but it is true for real-life survival data, especially. In typical survival time data sets one typically has a small effective sample size (number of observed events). We can easily imagine that for two different splits of the data, we can end up with quite different values for m_{stop} . It has been very effectively demonstrated that the split of the folds has a large impact on the choice of m_{stop} (Seibold et al., 2016). Seibold et al. (2016) suggest simply repeating the cross-validation scheme. They show that repeating even 5 times effectively averages out the randomness. In other words, we divide the data into K folds, and repeat this J times. Now let $\kappa(j, k)$ be the k 'th fold in the j 'th split. We end up with a new estimate for the prediction error,

$$\text{RCV}(m) = \sum_{j=1}^J \sum_{k=1}^K \sum_{i \in \kappa(j, k)} L((t_i, \delta_i), \theta_m; \mathbf{x}_i). \quad (3.6)$$

As before, we choose m_{stop} to be the minimizer of this error,

$$m_{\text{stop}} = \underset{m}{\operatorname{argmin}} \text{RCV}(m). \quad (3.7)$$

In practice, to ensure we find the minimizing m , we let the boosting algorithm run for $m = 1$ to $m = M$, where M is a large number that we are sure will result in a overfitted model.

derived from the pre-defined set of base learners. These base learners are usually relatively simple, regularized (see definition of regularization) parametric effects of β_j . Typical examples are such as linear least squares, stumps (trees with one split; see Bühlmann and Hothorn (2007) and Hastie et al. (2009)), or splines with a few degrees of freedom. In many cases, each base learner is defined on exactly one element x_j of x , and thus (3.2) simplifies to

$$\eta(x) = \beta_0 + \sum_{j=1}^p f_j(x_j | \beta_j). \quad (3.8)$$

{eq:eta-
componentwise}

To estimate the parameters β_1, \dots, β_j of the additive predictor, the boosting algorithm minimizes the empirical risk $\rho: \mathcal{Y} \times \mathbb{R} \rightarrow \mathbb{R}$, also called the in-sample

error and the training error, summed over all samples in the learning data set D ,

$$R(D) = \sum_{i=1}^n \rho(y^{(i)}, \eta(x^{(i)})). \quad (3.9)$$

{eq:empirical-risk}

$$\rho(y^{(i)}, \eta(x^{(i)})) = -l(y_0(x^{(i)}), \mu(x^{(i)}), t_i, d_i) \quad (3.10)$$

Let

$$y_0(x) = \beta^T \mathbf{x} = \beta_0 + \sum_{j=1}^p \beta_j x_j \quad (3.11)$$

and

$$\mu(z) = \gamma^T \mathbf{z} = \gamma_0 + \sum_{j=1}^p \gamma_j z_j \quad (3.12)$$

The goal is to estimate a function which minimizes the loss over an unseen “hold-out” sample, often called the out-of-sample error, the generalization error, or the test error.

$$\text{Err} = \rho(y, \eta(x)). \quad (3.13)$$

The main idea of boosting is to fit simple base learners $h(\cdot)$ one by one to the negative gradient vector of the loss $u = (u^{(1)}, u^{(2)}, \dots, u^{(n)})$, instead of to the true outcomes $y = (y^{(1)}, y^{(2)}, \dots, y^{(n)})$. Base learners are chosen in such a way that they approximate the effect $\hat{f}(x|\beta_j) = \sum_m h_j(\cdot)$. The negative gradient vector in iteration m , evaluated at the estimated additive predictor $\hat{\eta}^{[m-1]}(x^{(i)})$, is defined as

$$\mathbf{u} = \left(-\frac{\partial}{\partial \eta} \rho(y, \eta) \Big|_{\eta=\hat{\eta}^{[m-1]}(x^{(i)}), y=y^{(i)}} \right)_{i=1, \dots, n}.$$

In every boosting iteration, each base learner is fitted separately to the negative gradient vector by least-squares or penalized least-squares regression. The best fitting base learner is selected based on the residual sum of squares with respect to u ,

$$j^* = \underset{j \in \{1, \dots, J\}}{\text{argmin}} \sum_{i=1}^n (u^{(i)} - \hat{h}_j(x^{(i)}))^2.$$

Only the best performing base learner $\hat{h}_{j^*}(\cdot)$ will be used to update the current additive predictor, leading to the update being

$$\eta^{[m]}(\cdot) = \eta^{[m-1]}(\cdot) + \text{sl} \cdot \hat{h}_{j^*}(\cdot),$$

where $0 < \text{sl} < 1$ denotes the step length or learning rate. The choice of step length is not of critical importance as long as it is sufficiently small (Schmid and Hothorn, 2008), but the convention is to use $\text{sl} = 0.1$ (Mayr et al., 2014).

With a fixed step length (learning rate), the main tuning parameter for gradient boosting is the number of iterations m that are performed before the algorithm is stopped. We denote the resulting parameter m_{stop} . If m_{stop} is too small, the model will underfit and it cannot fully incorporate the influence of the effects on the response and will consequently have poor performance. On the other hand, too many iterations will result in overfitting, leading to poor generalization.

Chapter 4

First hitting time boost

In this chapter, we propose a component-wise boosting algorithm for fitting the inverse gaussian first hitting time model to survival data.

4.1 Algorithm

We apply the component-wise boosting algorithm 6 with loss function $\rho(\mu, \mathbf{y}_0) = -\log L(y_0, \mu)$. We differentiate the loss function with respect to these two and get For more details on the derivation, see A.

We might call this cyclical boosting.

Maybe use b instead of y_0 ,
to not get subscript chaos?

Boost in same

Another way to do this is to only boost one component in each iteration. The component might be corresponding to X , or it might be corresponding to Z .

algo: fhtboost

Algorithm 6 FHT Boost with twodimensional loss function

1. Initialize the n -dimensional vectors $\hat{y}_0^{[0]}, \hat{\mu}^{[0]}$, with offset values, e.g. with $\hat{y}_0^{[0]} = \mathbf{0}, \dots, \hat{\mu}^{[0]} = \mathbf{0}$. Alternatively, one can use the maximum likelihood estimates as offset values.
 2. For both components of the loss function, specify base learners, in particular, a component-wise base learner which can be used for each of the p variables used in \mathbf{X} corresponding to y_0 and the d variables in \mathbf{Z} corresponding to μ . Like earlier, the base learner takes one input variable and has one output variable. Examples include least squares linear regression.
 3. Set $m = 0$ and $\nu = 0.1$.
 4. Increase m by 1.
 - a) If $m > m_{\text{stop}, y_0}$, proceed to step 4 e). If not, compute the negative partial derivative $-\frac{\partial \rho}{\partial y_0}$ and evaluate at $\hat{f}^{[m-1]}(X_i, Z_i) = \left(\hat{y}_0^{[m-1]}(X_i), \hat{\mu}^{[m-1]}(Z_i) \right)_{i=1, \dots, n}$. This yields the negative gradient vector $U_{y_0}^{[m-1]} = \left(U_{i, y_0}^{[m-1]} \right)_{i=1, \dots, n} := \left(-\frac{\partial}{\partial y_0} \rho \left(Y_i, \hat{f}^{[m-1]}(X_i, Z_i) \right) \right)_{i=1, \dots, n}$.
 - b) Fit the negative gradient vector $U_{y_0}^{[m-1]}$ to each of the p components of \mathbf{X} separately (i.e. to each predictor variable) using the base learners specified in step 2. This yields p vectors of predicted values, where each vector is an estimate of the negative gradient vector $U_{y_0}^{[m-1]}$.
 - c) Select the component of \mathbf{X} which best fits $U_{y_0}^{[m-1]}$ according to R^2 . Set $\hat{U}_{y_0}^{[m-1]}$ equal to the fitted values of the corresponding best model fitted in the previous step.
 - d) Update $\hat{y}_0^{[m-1]} \leftarrow \hat{y}_0^{[m-1]} + \nu \hat{U}_{y_0}^{[m-1]}$.
 - e) If $m > m_{\text{stop}, \mu}$, proceed to step 4 j). If not, compute the negative partial derivative $-\frac{\partial \rho}{\partial \mu}$ and evaluate at $\hat{f}^{[m-1]}(X_i, Z_i) = \left(\hat{y}_0^{[m-1]}(X_i), \hat{\mu}^{[m-1]}(Z_i) \right)_{i=1, \dots, n}$. This yields the negative gradient vector $U_{\mu}^{[m-1]} = \left(U_{i, \mu}^{[m-1]} \right)_{i=1, \dots, n} := \left(-\frac{\partial}{\partial \mu} \rho \left(Y_i, \hat{f}^{[m-1]}(X_i, Z_i) \right) \right)_{i=1, \dots, n}$.
 - f) Fit the negative gradient vector $U_{\mu}^{[m-1]}$ to each of the p components of \mathbf{Z} separately (i.e. to each predictor variable) using the base learners specified in step 2. This yields d vectors of predicted values, where each vector is an estimate of the negative gradient vector $U_{\mu}^{[m-1]}$.
 - g) Select the component of \mathbf{Z} which best fits $U_{\mu}^{[m-1]}$ according to R^2 . Set $\hat{U}_{\mu}^{[m-1]}$ equal to the fitted values of the corresponding best model fitted in the previous step.
 - h) Update $\hat{\mu}^{[m-1]} \leftarrow \hat{\mu}^{[m-1]} + \nu \hat{U}_{\mu}^{[m-1]}$.
 - i) Update $\hat{f}^{[m]} \leftarrow \hat{f}^{[m-1]}$.
 - j) If $m > \max(m_{\text{stop}, y_0}, m_{\text{stop}, \mu})$, go to step 5. If not, repeat step 4.
 5. Return $\hat{f}^{[m]}$.
-

Appendices

Appendix A

Appendix 1: Differentiating the IG FHT

appendix

First we have the likelihood,

$$L(y_0, \mu) = \prod_{i=1}^n \left(\frac{y_0}{\sqrt{2\pi\sigma^2 t_i^3}} \exp \left[-\frac{(y_0 + \mu t_i)^2}{2\sigma^2 t_i} \right] \right)^{\delta_i} \times \left[1 - \Phi \left(-\frac{y_0 + \mu t_i}{\sqrt{\sigma^2 t_i}} \right) - \exp \left(-\frac{2y_0\mu}{\sigma^2} \right) \Phi \left(\frac{\mu t_i - y_0}{\sqrt{\sigma^2 t_i}} \right) \right]^{1-\delta_i}, \quad (\text{A.1})$$

with respect to parameters μ , and y_0 . First, note that for any cumulative distribution function F that is symmetric around 0, and for $x \in \mathbb{R}$,

$$F(x) = 1 - (1 - F(x)) = 1 - F(-x), \quad (\text{A.2})$$

and so in particular,

$$\Phi(x) = 1 - (1 - \Phi(x)) = 1 - \Phi(-x), \quad (\text{A.3})$$

and thus we can rewrite (A.1) as

$$L(y_0, \mu) = \prod_{i=1}^n \left(\frac{y_0}{\sqrt{2\pi\sigma^2 t_i^3}} \exp \left[-\frac{(y_0 + \mu t_i)^2}{2\sigma^2 t_i} \right] \right)^{\delta_i} \times \left[\Phi \left(\frac{y_0 + \mu t_i}{\sqrt{\sigma^2 t_i}} \right) - \exp \left(-\frac{2y_0\mu}{\sigma^2} \right) \Phi \left(\frac{\mu t_i - y_0}{\sqrt{\sigma^2 t_i}} \right) \right]^{1-\delta_i}. \quad (\text{A.4})$$

It is easier to work with the log likelihood, so we take the log of (A.4) and get

$$l(y_0, \mu) = \sum_{i=1}^n \delta_i \left(\ln y_0 - \frac{1}{2} \ln(2\pi\sigma^2 t_i^3) - \frac{(y_0 + \mu t_i)^2}{2\sigma^2 t_i} \right) + (1 - \delta_i) \ln \left(\Phi \left(\frac{\mu t_i + y_0}{\sqrt{\sigma^2 t_i}} \right) - \exp \left(-\frac{2y_0\mu}{\sigma^2} \right) \Phi \left(\frac{\mu t_i - y_0}{\sqrt{\sigma^2 t_i}} \right) \right) \quad (\text{A.5})$$

To make things easier, let us introduce some intermediate functions here. Let

$$\ln f_i(y_0, \mu) = \ln y_0 - \frac{1}{2} \ln(2\pi\sigma^2 t_i^3) - \frac{(y_0 + \mu t_i)^2}{2\sigma^2 t_i} \quad (\text{A.6})$$

and

$$S_i(y_0, \mu) = \Phi\left(\frac{\mu t_i + y_0}{\sqrt{\sigma^2 t_i}}\right) - \exp\left(-\frac{2y_0\mu}{\sigma^2}\right) \Phi\left(\frac{\mu t_i - y_0}{\sqrt{\sigma^2 t_i}}\right). \quad (\text{A.7})$$

So we get

$$l(y_0, \mu) = \sum_{i=1}^n \delta_i \ln f_i(y_0, \mu) + (1 - \delta_i) \ln S_i(y_0, \mu) \quad (\text{A.8})$$

Thus we see that the partial derivatives are

$$\frac{\partial}{\partial y_0} l(y_0, \mu) = \sum_{i=1}^n \delta_i \frac{\partial}{\partial y_0} \ln f_i(y_0, \mu) + (1 - \delta_i) \frac{\frac{\partial}{\partial y_0} S_i(y_0, \mu)}{S_i(\theta)} \quad (\text{A.9})$$

and

$$\frac{\partial}{\partial \mu} l(y_0, \mu) = \sum_{i=1}^n \delta_i \frac{\partial}{\partial \mu} \ln f_i(y_0, \mu) + (1 - \delta_i) \frac{\frac{\partial}{\partial \mu} S_i(y_0, \mu)}{S_i(\theta)} \quad (\text{A.10})$$

We take these one by one

$$\frac{\partial}{\partial y_0} \ln f_i(y_0, \mu) = \frac{1}{y_0} - \frac{y_0 + \mu t_i}{\sigma^2 t_i} \quad (\text{A.11})$$

$$\frac{\partial}{\partial \mu} \ln f_i(y_0, \mu) = -\frac{y_0 + \mu t_i}{\sigma^2} \quad (\text{A.12})$$

$$\begin{aligned} \frac{\partial}{\partial y_0} S_i(y_0, \mu) &= \frac{1}{\sqrt{\sigma^2 t_i}} \phi\left(\frac{\mu t_i + y_0}{\sqrt{\sigma^2 t_i}}\right) + \frac{2\mu}{\sigma^2} \exp\left(-\frac{2y_0\mu}{\sigma^2}\right) \Phi\left(\frac{\mu t_i - y_0}{\sqrt{\sigma^2 t_i}}\right) \\ &\quad + \frac{1}{\sqrt{\sigma^2 t_i}} \exp\left(-\frac{2y_0\mu}{\sigma^2}\right) \phi\left(\frac{\mu t_i - y_0}{\sqrt{\sigma^2 t_i}}\right) \end{aligned} \quad (\text{A.13})$$

$$\begin{aligned} \frac{\partial}{\partial \mu} S_i(y_0, \mu) &= \frac{t_i}{\sqrt{\sigma^2 t_i}} \phi\left(\frac{\mu t_i + y_0}{\sqrt{\sigma^2 t_i}}\right) + \frac{2y_0}{\sigma^2} \exp\left(-\frac{2y_0\mu}{\sigma^2}\right) \Phi\left(\frac{\mu t_i - y_0}{\sqrt{\sigma^2 t_i}}\right) \\ &\quad - \frac{t_i}{\sqrt{\sigma^2 t_i}} \exp\left(-\frac{2y_0\mu}{\sigma^2}\right) \phi\left(\frac{\mu t_i - y_0}{\sqrt{\sigma^2 t_i}}\right) \end{aligned} \quad (\text{A.14})$$

Hence

$$\begin{aligned} \frac{\partial}{\partial y_0} l(y_0, \mu) &= \sum_{i=1}^n \left(\delta_i \left(\frac{1}{y_0} - \frac{y_0 + \mu t_i}{\sigma^2 t_i} \right) + (1 - \delta_i) \left[\frac{1}{\sqrt{\sigma^2 t_i}} \phi\left(\frac{\mu t_i + y_0}{\sqrt{\sigma^2 t_i}}\right) + \frac{2\mu}{\sigma^2} \exp\left(-\frac{2y_0\mu}{\sigma^2}\right) \Phi\left(\frac{\mu t_i - y_0}{\sqrt{\sigma^2 t_i}}\right) \right. \right. \\ &\quad \left. \left. + \frac{1}{\sqrt{\sigma^2 t_i}} \exp\left(-\frac{2y_0\mu}{\sigma^2}\right) \phi\left(\frac{\mu t_i - y_0}{\sqrt{\sigma^2 t_i}}\right) \right] \left[\Phi\left(\frac{\mu t_i + y_0}{\sqrt{\sigma^2 t_i}}\right) - \exp\left(-\frac{2y_0\mu}{\sigma^2}\right) \Phi\left(\frac{\mu t_i - y_0}{\sqrt{\sigma^2 t_i}}\right) \right]^{-1} \right) \end{aligned} \quad (\text{A.15})$$

and

$$\begin{aligned}
& \frac{\partial}{\partial \mu} l(y_0, \mu) \\
&= \sum_{i=1}^n \left(\delta_i \left(-\frac{y_0 + \mu t_i}{\sigma^2} \right) + (1 - \delta_i) \left[\frac{t_i}{\sqrt{\sigma^2 t_i}} \phi \left(\frac{\mu t_i + y_0}{\sqrt{\sigma^2 t_i}} \right) + \frac{2y_0}{\sigma^2} \exp \left(-\frac{2y_0 \mu}{\sigma^2} \right) \Phi \left(\frac{\mu t_i - y_0}{\sqrt{\sigma^2 t_i}} \right) \right. \right. \\
&\quad \left. \left. - \frac{t_i}{\sqrt{\sigma^2 t_i}} \exp \left(-\frac{2y_0 \mu}{\sigma^2} \right) \phi \left(\frac{\mu t_i - y_0}{\sqrt{\sigma^2 t_i}} \right) \right] \left[\Phi \left(\frac{\mu t_i + y_0}{\sqrt{\sigma^2 t_i}} \right) - \exp \left(-\frac{2y_0 \mu}{\sigma^2} \right) \Phi \left(\frac{\mu t_i - y_0}{\sqrt{\sigma^2 t_i}} \right) \right]^{-1} \right) \\
&\hspace{15em} \text{(A.16)}
\end{aligned}$$

Bibliography

ABG	Aalen, O., Borgan, O., and Gjessing, H. (2008). <i>Survival and Event History Analysis: A Process Point of View</i> . Statistics for Biology and Health. Springer New York.
bauer-kohavi	Bauer, E. and Kohavi, R. (1999). An empirical comparison of voting classification algorithms: Bagging, boosting, and variants. <i>Machine Learning</i> , 36(1):105–139.
breiman1998	Breiman, L. (1998). Arcing classifier (with discussion and a rejoinder by the author). <i>Ann. Statist.</i> , 26(3):801–849.
bovelstadborgan	Bøvelstad, H. M. and Borgan, Ø. (2011). Assessment of evaluation criteria for survival prediction from genomic data. <i>Biometrical Journal</i> , 53(2):202–216.
buhlmann2006	Bühlmann, P. (2006). Boosting for high-dimensional linear models. <i>Ann. Statist.</i> , 34(2):559–583.
buhlmann2007	Bühlmann, P. and Hothorn, T. (2007). Boosting algorithms: Regularization, prediction and model fitting. <i>Statist. Sci.</i> , 22(4):477–505.
buhlmann-yu	Bühlmann, P. and Yu, B. (2003). Boosting with the l2 loss. <i>Journal of the American Statistical Association</i> , 98(462):324–339.
caroni2017	Caroni, C. (2017). <i>First Hitting Time Regression Models</i> . John Wiley & Sons, Inc.
chang2010	Chang, Y.-C. I., Huang, Y., and Huang, Y.-P. (2010). Early stopping in l2boosting. <i>Computational Statistics & Data Analysis</i> , 54(10):2203 – 2213.
chhikara1988	Chhikara, R. (1988). <i>The Inverse Gaussian Distribution: Theory: Methodology, and Applications</i> . Statistics: A Series of Textbooks and Monographs. Taylor & Francis.
cox1965	Cox, D. and Miller, H. (1965). <i>The theory of stochastic processes</i> . Wiley publications in statistics. Wiley.
cox	Cox, D. R. (1992). <i>Regression Models and Life-Tables</i> , pages 527–541. Springer New York, New York, NY.
adaboost	Freund, Y. and Schapire, R. E. (1996). Experiments with a new boosting algorithm. In <i>Proceedings of the Thirteenth International Conference on International Conference on Machine Learning</i> , ICML’96, pages 148–156, San Francisco, CA, USA. Morgan Kaufmann Publishers Inc.

- | |
|--------------|
| friedman2000 |
|--------------|
- Friedman, J., Hastie, T., and Tibshirani, R. (2000). Additive logistic regression: a statistical view of boosting (with discussion and a rejoinder by the authors). *Ann. Statist.*, 28(2):337–407.
- | |
|--------------|
| friedman2001 |
|--------------|
- Friedman, J. H. (2001). Greedy function approximation: A gradient boosting machine. *Ann. Statist.*, 29(5):1189–1232.
- | |
|------------|
| hastie2007 |
|------------|
- Hastie, T. (2007). Comment: Boosting algorithms: Regularization, prediction and model fitting. *Statist. Sci.*, 22(4):513–515.
- | |
|-----|
| ESL |
|-----|
- Hastie, T., Tibshirani, R., and Friedman, J. (2009). *The Elements of Statistical Learning: Data Mining, Inference, and Prediction*. Springer series in statistics. Springer.
- | |
|---------------|
| kearnsvaliant |
|---------------|
- Kearns, M. and Valiant, L. G. (1989). Cryptographic limitations on learning boolean formulae and finite automata. In *Proceedings of the Twenty-first Annual ACM Symposium on Theory of Computing*, STOC '89, pages 433–444, New York, NY, USA. ACM.
- | |
|--------|
| kohavi |
|--------|
- Kohavi, R. (1995). A study of cross-validation and bootstrap for accuracy estimation and model selection. In *Proceedings of the 14th International Joint Conference on Artificial Intelligence - Volume 2*, IJCAI'95, pages 1137–1143, San Francisco, CA, USA. Morgan Kaufmann Publishers Inc.
- | |
|-------------|
| lachenbruch |
|-------------|
- Lachenbruch, P. A. and Mickey, M. R. (1968). Estimation of error rates in discriminant analysis. *Technometrics*, 10(1):1–11.
- | |
|-----------------|
| leewhitmore2006 |
|-----------------|
- Lee, M.-L. T. and Whitmore, G. A. (2006). Threshold regression for survival analysis: Modeling event times by a stochastic process reaching a boundary. *Statist. Sci.*, 21(4):501–513.
- | |
|---------|
| mayr14a |
|---------|
- Mayr, A., Binder, H., Gefeller, O., and Schmid, M. (2014). The evolution of boosting algorithms. from machine learning to statistical modelling. *Methods of Information in Medicine*, 53(6):419–427.
- | |
|-------------------|
| gamboostlss-paper |
|-------------------|
- Mayr, A., Fenske, N., Hofner, B., Kneib, T., and Schmid, M. (2012a). Generalized additive models for location, scale and shape for high dimensional data—a flexible approach based on boosting. *Journal of the Royal Statistical Society. Series C (Applied Statistics)*, 61(3):403–427.
- | |
|-------------|
| mayr-hofner |
|-------------|
- Mayr, A., Hofner, B., and Schmid, M. (2012b). The importance of knowing when to stop. a sequential stopping rule for component-wise gradient boosting. *Methods of Information in Medicine*, 51(2):178–186.
- | |
|-------|
| Rlang |
|-------|
- R Core Team (2013). *R: A Language and Environment for Statistical Computing*. R Foundation for Statistical Computing, Vienna, Austria.
- | |
|--------|
| gamlss |
|--------|
- Rigby, R. A. and Stasinopoulos, D. M. (2005). Generalized additive models for location, scale and shape. *Journal of the Royal Statistical Society. Series C (Applied Statistics)*, 54(3):507–554.
- | |
|----------------|
| schmid-hothorn |
|----------------|
- Schmid, M. and Hothorn, T. (2008). Boosting additive models using component-wise p-splines. *Comput. Stat. Data Anal.*, 53(2):298–311.

schmid

Schmid, M., Potapov, S., Pfahlberg, A., and Hothorn, T. (2010). Estimation and regularization techniques for regression models with multidimensional prediction functions. *Statistics and Computing*, 20(2):139–150.

seibold

Seibold, H., Bernau, C., Boulesteix, A.-L., and Bin, R. D. (2016). On the choice and influence of the number of boosting steps.

thomas2018

Thomas, J., Mayr, A., Bischl, B., Schmid, M., Smith, A., and Hofner, B. (2018). Gradient boosting for distributional regression: faster tuning and improved variable selection via noncyclical updates. *Statistics and Computing*, 28(3):673–687.

tukey

Tukey, J. (1977). *Exploratory Data Analysis*. Addison-Wesley series in behavioral science. Addison-Wesley Publishing Company.

whitmore1986

Whitmore, G. A. (1986). First-passage-time models for duration data: Regression structures and competing risks. *Journal of the Royal Statistical Society. Series D (The Statistician)*, 35(2):207–219.

threg

Xiao, T., Whitmore, G., He, X., and Lee, M.-L. (2015). The r package threg to implement threshold regression models. *Journal of Statistical Software, Articles*, 66(8):1–16.