

ML Preparation Notes

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Chapter 1

Basic Statistics

In this section, we review some basic concepts from Statistics. The material in this section is based on [4].

1.1 Covariance and Correlation

The *correlation* between two sets of data is a measure of the strength of the relationship between them. In particular, Pearson's correlation coefficient is a measure of linear relationship between two sets of data. Let X and Y be two random variables. Then Pearson's correlation coefficient $\rho(X, Y)$ is defined as:

$$\rho(X, Y) = \frac{\text{Cov}(X, Y)}{\sigma_X \sigma_Y} \quad (1.1)$$

Two important facts about the Pearson's correlation coefficient [3]:

1. $-1 \leq \rho(X, Y) \leq 1$
2. $|\rho(X, Y)| = 1$ iff there exists $a \neq 0$ and b such that $Y = aX + b$.

1.2 Tail Inequalities and the Law of Large Numbers

We first look at Markov and Chebyshev inequalities and then apply them to random samples.

The Markov inequality tells us how much probability mass can be at large values given the mean of the distribution.

Theorem 1.1. *Let X be a random variable that takes on positive values only. Then for $t > 0$,*

$$P(X \geq t) \leq \frac{E[X]}{t}. \quad (1.2)$$

Proof. Let us assume that X is a continuous r.v. with pdf f_X . By definition,

$$\begin{aligned} E[X] &= \int_0^\infty x f_X(x) dx \\ &= \int_0^t x f_X(x) dx + \int_t^\infty x f_X(x) dx \\ &\geq t \int_t^\infty f_X(x) dx. \end{aligned}$$

Since $t > 0$, dividing both sides of the last inequality yields the Markov inequality. □

This inequality is useful when $t > E[X]$; when $t \leq E[X]$, it merely bounds $P(X \geq t)$ by 1.

The Chebyshev inequality involves both the mean and the variance of the distribution. It bounds the probability of how far a random variable can be from its mean as a function of the variance.

Theorem 1.2. *Let X be a random variable for which $\text{Var}(X)$ exists. Then*

$$P(|X - E[X]| > t) \leq \frac{\text{Var}(X)}{t^2}. \quad (1.3)$$

Proof. Define $Y := (X - E[X])^2$ so that $E[Y] = \text{Var}(X)$. Applying the Markov inequality to the r.v. Y , we obtain:

$$P(Y \geq t^2) \leq \frac{\text{Var}(X)}{t^2}.$$

But $P(Y \geq t^2) = P(|X - E[X]| > t)$ and so this proves the Chebyshev inequality too. \square

To talk about the law of large numbers, one has to talk about the notion of a sequence of random variables converging to a real number. We say that a sequence Z_1, Z_2, \dots of random variables converges to the number b if the distribution of Z_n as $n \rightarrow \infty$ becomes more and more concentrated around this single number.

Definition 1.1 (Convergence in Probability). A sequence Z_1, Z_2, \dots of random variables converges in probability to the number b if for every $\epsilon > 0$, the following condition holds:

$$\lim_{n \rightarrow \infty} P(|Z_n - b| < \epsilon) = 1. \quad (1.4)$$

This fact is denoted by $Z_n \xrightarrow{p} b$.

Theorem 1.3 (The Law of Large Numbers). *Let X_1, \dots, X_n be a random sample from a distribution with mean μ and finite variance σ^2 . Then the sample mean \bar{X}_n converges in probability to μ .*

Proof. We have $E[\bar{X}_n] = \mu$ and $\text{Var}(\bar{X}_n) = \sigma^2/n$. Use the Chebyshev inequality to obtain:

$$P(|\bar{X}_n - \mu| \leq t) \geq 1 - \frac{\sigma^2}{nt^2}.$$

For every fixed $t > 0$, take limits as $n \rightarrow \infty$ to obtain: $\lim_{n \rightarrow \infty} P(|\bar{X}_n - \mu| \leq t) = 1$. Hence $\bar{X} \xrightarrow{p} \mu$. \square

If we know that a sequence of random variables converges in probability, what can we say about continuous functions of that sequence of random variables?

Theorem 1.4 (Continuous Functions of Random Variables). *Suppose a sequence Z_1, Z_2, \dots converges in probability to b and that g is a function continuous at b . Then $g(Z_n) \xrightarrow{p} g(b)$.*

Proof. We have to show $\lim_{n \rightarrow \infty} P(|g(Z_n) - g(b)| < \epsilon) = 1$. Since g is continuous at b , given $\epsilon > 0$ there exists $\delta > 0$ such that $|z_n - b| < \delta$ implies $|g(z_n) - g(b)| < \epsilon$. Now consider the events $\mathcal{E}_1 = \{|Z_n - b| < \delta\}$ and $\mathcal{E}_2 = \{|g(Z_n) - g(b)| < \epsilon\}$. By the continuity of g , \mathcal{E}_1 implies \mathcal{E}_2 , and we have that $P(\mathcal{E}_2) \geq P(\mathcal{E}_1)$. Since $\lim_{n \rightarrow \infty} P(\mathcal{E}_1) = 1$, we must have $\lim_{n \rightarrow \infty} P(\mathcal{E}_2) = 1$. \square

Theorem 1.5 (Chernoff Bounds). *Let X be a random variable with moment generating function ψ_X . Then for any real t ,*

$$P(X \geq t) \leq \min_{s>0} e^{-st} \psi_X(s). \quad (1.5)$$

Proof. Fix $s > 0$. Since the map $x \rightarrow e^x$ is 1-1 and increasing, we have that $X \geq t$ if and only if $e^{sX} \geq e^{st}$. Therefore by the Markov inequality, we obtain:

$$P(X \geq t) = P(e^{sX} \geq e^{st}) \leq \frac{E[e^{sX}]}{e^{st}} = e^{-st} \psi_X(s).$$

This holds for any $s > 0$ and, in particular, when we minimize $e^{-st} \psi_X(s)$ subject to the condition that $s > 0$. \square

Example 1.1. For each positive integer n , let X_n be a nonnegative random variable with finite mean μ_n . Show that if $\lim_{n \rightarrow \infty} \mu_n = 0$, then $X_n \xrightarrow{p} 0$.

Solution. We have to show that for every $\epsilon > 0$, $\lim_{n \rightarrow \infty} P(X_n \geq \epsilon) = 0$. To this end, fix $\epsilon > 0$ and $\delta > 0$. Since $\lim_{n \rightarrow \infty} \mu_n = 0$, there exists $n_{\epsilon\delta} \in \mathbf{N}$ such that for all $n \geq n_{\epsilon\delta}$, $\mu_n < \epsilon \cdot \delta$. Combining this with the Markov inequality, we obtain that for all $n \geq n_{\epsilon\delta}$

$$P(X_n \geq \epsilon) \leq \frac{\mu_n}{\epsilon} < \delta.$$

This is precisely what we need to show in order to prove that $\lim_{n \rightarrow \infty} P(X_n \geq \epsilon) = 0$. \blacksquare

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Example 1.2. Let Z_1, Z_2, \dots be a sequence of random variables, and suppose that, for $n = 1, 2, \dots$, the distribution of Z_n is as follows:

$$P(Z_n = n^2) = \frac{1}{n} \text{ and } P(Z_n = 0) = 1 - \frac{1}{n}.$$

Show that $\lim_{n \rightarrow \infty} E[Z_n] = \infty$ but $Z_n \xrightarrow{p} 0$.

Solution. Now $E[Z_n] = n^2 \cdot \frac{1}{n} = n$ and clearly $E[Z_n] \rightarrow \infty$ as $n \rightarrow \infty$. Since $\lim_{n \rightarrow \infty} P(Z_n = 0) = 1$, it follows that for any $\epsilon > 0$, $\lim_{n \rightarrow \infty} P(Z_n > \epsilon) = 0$. This shows that $Z_n \xrightarrow{p} 0$. \blacksquare

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1.3 The Central Limit Theorem

The CLT is a formal statement of how normal distributions can approximate distributions of general sums or averages of iid random variables. To start with, consider the case when X_1, \dots, X_n are iid random variables with $X_i \sim N(\mu_i, \sigma_i^2)$ for $1 \leq i \leq n$. Let $Y = a_1 X_1 + \dots + a_n X_n + b$, where $a_i \neq 0$ for at least one $1 \leq i \leq n$. Then Y is normally distributed as the following shows.

Theorem 1.6. Let X_1, \dots, X_n be iid normally distributed random variables with X_i having mean μ_i and variance σ_i^2 , for $1 \leq i \leq n$. Define $Y = a_1 X_1 + \dots + a_n X_n + b$, where $a_i \neq 0$ for at least one $1 \leq i \leq n$. Then Y is normally distributed with mean $a_1 \mu_1 + \dots + a_n \mu_n + b$ and variance $a_1^2 \mu_1^2 + \dots + a_n^2 \mu_n^2$.

Proof. Let $\psi_i(t)$ be the mgf of X_i for $1 \leq i \leq n$, and let $\psi(t)$ denote the mgf of $a_1X_1 + \cdots + a_nX_n + b$. Then

$$\begin{aligned}\psi(t) &= \mathbb{E}[e^{t(a_1X_1 + \cdots + a_nX_n + b)}] \\ &= e^{tb} \cdot \prod_{i=1}^n \mathbb{E}[e^{a_iX_it}] \\ &= e^{tb} \cdot \prod_{i=1}^n \psi_i(a_it) \\ &= e^{tb} \cdot \prod_{i=1}^n \exp\left(a_i\mu_it + \frac{1}{2}a_i^2\sigma_i^2t^2\right) \\ &= \exp\left(\left(b + \sum_i \mu_i\right)t + \frac{1}{2}\left(\sum_i a_i^2\sigma_i^2\right)t^2\right).\end{aligned}$$

Note that $\psi(t)$ is the mgf of a normal distribution with mean $b + \sum_i \mu_i$ and variance $\sum_i a_i^2\sigma_i^2$. Hence $Y \sim N(b + \sum_i \mu_i, \sum_i a_i^2\sigma_i^2)$. \square

Theorem 1.7 (Central Limit Theorem for IID RVs). *Let X_1, \dots, X_n be iid random variables from a distribution with mean μ and variance $\sigma^2 < \infty$. Then for each fixed number x ,*

$$\lim_{n \rightarrow \infty} \mathbb{P}\left(\frac{\bar{X}_n - \mu}{\sigma/\sqrt{n}} \leq x\right) = \Phi(x), \quad (1.6)$$

where Φ denotes the cdf of the standard normal distribution.

We can interpret Equation 1.6 as follows: If we have a large random sample X_1, \dots, X_n from an arbitrary distribution (whether discrete or continuous), the random variable $(\bar{X}_n - \mu)/(\sigma/\sqrt{n})$ is distributed approximately as the standard normal. Consequently, \bar{X}_n is distributed approximately as $N(\mu, \sigma^2/n)$ and $\sum_{i=1}^n X_i$ is distributed approximately as $N(n\mu, n\sigma^2)$.

1.4 Maximum Likelihood Estimation

Let the random variables X_1, \dots, X_n form a random sample from a distribution with pdf $f(x | \theta)$. Recall that this means that $X_i \stackrel{\text{iid}}{\sim} f(\cdot | \theta)$ for all $1 \leq i \leq n$. Let $f_n(\vec{x} | \theta)$ denote the value of the joint pdf of the random vector $(X_1, \dots, X_n)'$ at the point $\vec{x} = (x_1, \dots, x_n)'$. The *likelihood function* is the joint pdf of the observations of a random sample viewed as a function of θ for a given set of values of the sample. The maximum likelihood estimate of θ is that value of θ for which $f_n(\vec{x} | \theta)$ is maximized.

Example 1.3. Suppose that X_1, \dots, X_n form a random sample from a distribution with pdf $f(x | \theta)$ defined as follows:

$$f(x | \theta) = \begin{cases} \theta x^{\theta-1} & \text{for } 0 < x < 1 \\ 0 & \text{otherwise.} \end{cases}$$

Let us assume that $\theta > 0$. The joint distribution of the vector $(X_1, \dots, X_n)'$ at the point $\vec{x} = (x_1, \dots, x_n)'$ is

$$f_n(\vec{x} | \theta) = \prod_{i=1}^n \theta x_i^{\theta-1},$$

where we assume that $0 < x_i < 1$ for all i . Taking logs, we obtain that $\log f_n(\vec{x} \mid \theta) = n \log \theta + (\theta - 1) \sum_i \log x_i$. Take the derivative wrt θ and set to 0:

$$\frac{\partial \log f_n(\vec{x} \mid \theta)}{\partial \theta} = \frac{n}{\theta} + \sum_i \log x_i \stackrel{\text{set}}{=} 0$$

to obtain that $\theta = n / \sum_i \log \frac{1}{x_i}$. ▷

Example 1.4. Suppose that X_1, \dots, X_n form a random sample from a distribution with pdf $f(x \mid \theta)$ defined as follows:

$$f(x \mid \theta) = \frac{1}{2} e^{-|x-\theta|} \quad \text{for } -\infty < x < \infty.$$

Suppose that θ is unknown and that $-\infty < \theta < \infty$. In this case, the joint probability distribution is easily seen to be:

$$f_n(\vec{x} \mid \theta) = \frac{1}{2^n} e^{-\sum_i |x_i - \theta|}.$$

Take logs to obtain: $\log f_n(\vec{x} \mid \theta) = \log \frac{1}{2^n} - \sum_i |x_i - \theta|$. Maximizing f_n is equivalent to minimizing $\sum_i |x_i - \theta|$. This is equivalent to obtaining a point on the real line that minimizes the sum of the distances to the points x_1, \dots, x_n . This happens when θ is the median of x_1, \dots, x_n . ▷

Example 1.5. Suppose that X_1, \dots, X_n form a random sample from the uniform distribution on the interval $[\theta_1, \theta_2]$, where both θ_1 and θ_2 are unknown ($-\infty < \theta_1 < \theta_2 < \infty$). In this case, the log pdf of the joint distribution of $(X_1, \dots, X_n)'$ is given by

$$\log f_n(\vec{x} \mid \theta_1, \theta_2) = \log \prod_{i=1}^n \frac{1}{\theta_2 - \theta_1} = -n \log(\theta_2 - \theta_1).$$

Maximizing the likelihood is equivalent to minimizing $\log(\theta_2 - \theta_1)$. The minimum possible value of θ_2 is $\max\{x_1, \dots, x_n\}$ and the maximum possible value of θ_1 is $\min\{x_1, \dots, x_n\}$. ▷

Example 1.6. Suppose that a certain large population contains k different types of individuals ($k \geq 2$), and let θ_i denote the proportion of people of type i , for $1 \leq i \leq k$. Here, $0 \leq \theta_i \leq 1$ and $\sum_{i=1}^k \theta_i = 1$. Suppose also that in a random sample of n individuals from this population there are exactly n_i individuals of type i so that $n = n_1 + \dots + n_k$.

In this setting, for $1 \leq i \leq k$, define X_i to be the number of individuals of type i in a random sample of size n . Then the probability that $\bigwedge_{i=1}^k X_i = n_i$ is given by $\theta_1^{n_1} \dots \theta_k^{n_k}$. The log pdf of the joint distribution is given by:

$$\log f_k((n_1, \dots, n_k) \mid \theta_1, \dots, \theta_k, n) = \sum_{i=1}^k n_i \log \theta_i.$$

Note that there are actually $k - 1$ variables here since we may write $\theta_k = 1 - \sum_{i=1}^{k-1} \theta_i$. Differentiating wrt θ_i for $1 \leq i \leq k - 1$, we obtain:

$$\frac{\partial \log f_k}{\partial \theta_i} = \frac{n_i}{\theta_i} - \frac{n_k}{\theta_k}.$$

Setting this to 0, we get that $\theta_i / \theta_k = n_i / n_k$. Sum this up from $1 \leq i \leq k - 1$, to obtain:

$$\frac{1 - \theta_k}{\theta_k} = \frac{n - n_k}{n_k},$$

which yields $\theta_k = n_k / n$. Substitute this in $\theta_i / \theta_k = n_i / n_k$ to obtain $\theta_i = n_i / n$. ▷

Example 1.7 (Nonexistence of an MLE). An obvious disadvantage of the technique of maximum likelihood estimation is when the maximum does not exist. Consider again Example 1.5 where we let $\theta_1 = 0$ and $\theta = \theta_2$. The pdf of the uniform distribution is defined as:

$$f(x | \theta) = \begin{cases} \frac{1}{\theta} & 0 \leq x \leq \theta \\ 0 & \text{otherwise.} \end{cases}$$

Let us modify the definition of the pdf so that we use strict inequalities $0 < x < \theta$ above. Given a sample x_1, \dots, x_n , the log pdf is $-n \log \theta$ as before and the MLE technique would require us to minimize $\log \theta$. However, in this case, there is no $\theta > \max\{x_1, \dots, x_n\}$ that minimizes $\log \theta$ and the MLE does not exist. \triangleright

Example 1.8 (Non-uniqueness of an MLE). Consider a random sample X_1, \dots, X_n from the uniform distribution over the interval $[\theta, \theta + 1]$. In this case, the joint pdf is given by:

$$f_n(\vec{x} | \theta) = \begin{cases} 1 & \text{for } \theta \leq x_i \leq \theta + 1 \quad (1 \leq i \leq n) \\ 0 & \text{otherwise.} \end{cases}$$

In this case, the condition $\theta \leq x_i \leq \theta + 1$ for $1 \leq i \leq n$ may be written using the two conditions:

$$\theta \leq \min\{x_1, \dots, x_n\} \text{ and } \max\{x_1, \dots, x_n\} - 1 \leq \theta.$$

Any value of θ in the interval $[\max\{x_1, \dots, x_n\} - 1, \min\{x_1, \dots, x_n\}]$ is valid, but there is no unique value of θ . \triangleright

1.5 Bayesian Statistics and MCMC

This section is based on Chapters 12–15 from [6]. Bayes' rule gives us a recipe for calculating the posterior probability density.

$$P(\Theta | \text{data}) = \frac{P(\text{data} | \Theta) \cdot P(\Theta)}{P(\text{data})}. \quad (1.7)$$

Consider a case in which we have a sample of N data points x_1, \dots, x_N . We assume that the likelihood is a Poisson distribution with mean λ and that the prior for λ is a log-normal(1, 1) distribution. To calculate the probability of the data $P(\text{data})$, we must evaluate the integral:

$$P(\text{data}) = \int_0^\infty \prod_{i=1}^N \frac{\lambda^{x_i} e^{-\lambda}}{x_i!} \cdot \frac{1}{\sqrt{2\pi}\lambda} e^{-\frac{1}{2}(\log \lambda - 1)^2} d\lambda. \quad (1.8)$$

While this integral is not too difficult, it explains the problem of calculating posteriors analytically. As the number of parameters (the length of Θ) increases, calculating the probability of the data requires evaluating integrals in higher dimensional spaces. This is why we use alternative methods to derive approximate versions of the posterior.

In certain special cases, the posterior distribution can be easily derived. This occurs, for example, when the prior distribution is from a so-called *conjugate prior* family. A conjugate prior family is a set of distributions defined wrt a likelihood function. If one chooses the prior to be from this family, then the posterior is guaranteed to be from the same family. In such cases, one can dispense with computing integrals to find out the posterior.

Example 1.9 (Beta-Binomial). Let $X_1, \dots, X_n \stackrel{\text{iid}}{\sim} \text{Bernoulli}(\theta)$ and suppose that $\theta \sim \text{Beta}(\alpha, \beta)$. Suppose that one observes $X_i = x_i$ for $1 \leq i \leq n$. Then the posterior distribution of θ is given by:

$$f(\theta \mid X_1, \dots, X_n) = \frac{f(X_1, \dots, X_n \mid \theta) \cdot f(\theta)}{f(X_1, \dots, X_n)}.$$

In the above expression, $f(X_1, \dots, X_n)$ is a constant for a given set of values of the random variables X_i . The likelihood, when $X_i = x_i$ for $1 \leq i \leq n$, is given by:

$$\begin{aligned} f(x_1, \dots, x_n \mid \theta) &= \prod_{i=1}^n \theta^{x_i} (1 - \theta)^{1-x_i} \\ &= \theta^{\sum_i x_i} \cdot (1 - \theta)^{n - \sum_i x_i}. \end{aligned}$$

Now we may write:

$$\begin{aligned} f(\theta \mid X_1, \dots, X_n) &= \frac{f(X_1, \dots, X_n \mid \theta) \cdot f(\theta)}{f(X_1, \dots, X_n)} \\ &\propto \theta^{\sum_i x_i} \cdot (1 - \theta)^{n - \sum_i x_i} \cdot \theta^{\alpha-1} \cdot (1 - \theta)^{\beta-1} \\ &\propto \theta^{\alpha + \sum_i x_i - 1} \cdot (1 - \theta)^{\beta + n - \sum_i x_i - 1}. \end{aligned}$$

The last expression is the kernel of the beta distribution with parameters $\alpha + \sum_i x_i$ and $\beta + n - \sum_i x_i$. Since the expression on the right must be a valid probability distribution, the proportionality constant must be that of the beta distribution with these parameters, that is,

$$\frac{\Gamma(\alpha + \beta + n)}{\Gamma(\alpha + \sum_i x_i) \cdot \Gamma(\beta + n - \sum_i x_i)}.$$

▷

Thus if the prior is a beta distribution and the likelihood is a bernoulli distribution, then the posterior is a beta distribution also. Thus the set of beta distributions is a conjugate family wrt the bernoulli likelihood.

Chapter 2

Timeseries

The material in the section follows [2]. A *time series* is a sequence of random variables $\{X_1, X_2, \dots\}$. A complete model of such a time series would require specifying the joint probability distributions of all random vectors $(X_1, \dots, X_n)'$ for all $n \geq 1$. In practice, this might be impossible to do unless the time series is generated by some simple well-understood mechanism. Instead one typically specifies the first- and second-order moments of the joint distributions, that is, $E[X_t]$ and $E[X_t X_{t+h}]$ for all $t \geq 1$ and for all $h \geq 0$.

The next important concepts are those of stationarity and the auto-correlation function. Roughly speaking, a time series $\{X_t\}_{t=-\infty}^{\infty}$ is stationary if its “statistical properties” are similar to those of the time-shifted series $\{X_{t+h}\}_{t=-\infty}^{\infty}$ for every integer “lag” h . By statistical properties, we mean the first- and second-order moments of $\{X_t\}$.

Formally, a time series $\{X_t\}$ is weakly stationary if

1. $E[X_t]$ is independent of t
2. $\text{Cov}(X_t, X_{t+h})$ is independent of t for every fixed lag h .

In contrast, a time series $\{X_t\}$ is strictly stationary if the random vectors $(X_{t_1}, \dots, X_{t_n})'$ and $(X_{t_1+h}, \dots, X_{t_n+h})'$ have the same joint distributions for all sets of indices $\{t_1, \dots, t_n\}$, for all $h \geq 0$ and all $n \geq 1$. This is written as:

$$(X_{t_1}, \dots, X_{t_n})' \stackrel{d}{=} (X_{t_1+h}, \dots, X_{t_n+h})'$$

Strict stationarity implies the following:

1. The random variables X_t are identically distributed.
2. Pairs of random variables (X_t, X_{t+h}) have the same distribution as (X_1, X_{1+h}) (set $n = 2$).
3. Strict stationarity implies weak stationarity: $E[X_t] = E[X_1]$ and $\text{Cov}(X_t, X_{t+h}) = \text{Cov}(X_1, X_{1+h})$ for all $t \geq 1$ and all $h \geq 0$. Both terms are independent of t .
4. Weak stationarity does *not* imply strong stationarity. We show this by an example. Let $Z_i \stackrel{\text{iid}}{\sim} N(0, 1)$ for all i . Define X_t as:

$$X_t = \begin{cases} Z_t & \text{if } t \text{ is even} \\ 2Z_t & \text{if } t \text{ is odd.} \end{cases}$$

Then $E[X_t] = 0$ for all t . Also,

$$\text{Cov}(X_t, X_{t+h}) = \begin{cases} 0 & \text{if } h > 0 \\ 1 & \text{if } h = 0. \end{cases}$$

This follows from the assumed independence of X_t and X_{t+h} when $h > 0$. However, X_t s do not have the same distribution, a requirement of strict stationarity.

The autocovariance function of a stationary time series $\{X_t\}$ at lag h is defined as $\text{Cov}(X_{t+h}, X_t) = \text{Cov}(X_h, X_0)$. The autocorrelation of $\{X_t\}$ at lag h is defined as

$$\frac{\text{Cov}(X_{t+h}, X_t)}{\text{Var}(X_t)} = \frac{\text{Cov}(X_h, X_0)}{\text{Var}(X_0)}.$$

Example 2.1 (iid Noise). Suppose $\{X_t\}$ is iid noise with zero mean and $E[X_t^2] = \sigma^2 < \infty$. Then $E[X_t] = 0$ for all t and for every fixed lag h and all t :

$$\text{Cov}(X_t, X_{t+h}) = \begin{cases} 0 & \text{if } h > 0 \\ \sigma^2 & \text{if } h = 0. \end{cases}$$

Thus iid noise is stationary. We denote such a series as $\{X_t\} \sim \text{IID}(0, \sigma^2)$.

Example 2.2 (White Noise). A sequence $\{X_t\}$ of *uncorrelated* random variables with zero mean and finite second moment σ^2 is called *white noise*. Clearly, the covariance function at lag h is the same as that of iid noise and, as such, white noise is stationary. We denote such a series as $\{X_t\} \sim \text{WN}(0, \sigma^2)$. Unlike iid noise, the components of white noise need not be independent (recall that independence implies zero correlation but not the other way around). In particular, every $\text{IID}(0, \sigma^2)$ sequence is a $\text{WN}(0, \sigma^2)$ sequence but not the other way around.

Example 2.3 (Random Walk). A *random walk* $\{S_t\}$ is a sequence obtained by cumulatively summing iid random variables. A random walk with zero mean is obtained by defining $S_0 = 0$ and

$$S_t = X_1 + \cdots + X_t$$

for all $t > 0$, where $\{X_t\} \sim \text{IID}(0, \sigma^2)$. In this case, $E[S_t] = 0$ and $E[S_t^2] = t\sigma^2 < \infty$ for all t . For all lags $h \geq 0$ and all t ,

$$\begin{aligned} \text{Cov}(S_t, S_{t+h}) &= \text{Cov}(S_t, S_t + X_{t+1} + \cdots + X_{t+h}) \\ &= \text{Cov}(S_t, S_t) + \text{Cov}(S_t, X_{t+1}) + \cdots + \text{Cov}(S_t, X_{t+h}) \\ &= t\sigma^2. \end{aligned}$$

The last equality follows since $\text{Cov}(S_t, X_{t+i}) = 0$ for all $i \geq 1$. Hence $\text{Cov}(S_t, S_{t+h})$ depends on t and $\{S_t\}$ is not stationary.

Example 2.4 (First-Order Moving Average). Consider the series defined by

$$X_t = Z_t + \theta Z_{t-1}, \quad t = 0, \pm 1, \pm 2, \dots, \quad (2.1)$$

where $Z_t \sim \text{WN}(0, \sigma^2)$ and θ is a real-valued constant. Now $E[X_t] = 0$ and

$$\begin{aligned} E[X_t^2] &= E[(Z_t + \theta Z_{t-1})^2] \\ &= E[Z_t^2 + 2\theta Z_t Z_{t-1} + \theta^2 Z_{t-1}^2] \\ &= \sigma^2 + 2\theta E[Z_t Z_{t-1}] + \theta^2 \sigma^2 \\ &= (1 + \theta^2)\sigma^2. \end{aligned}$$

The last equality follows since, Z_t and Z_{t-1} being uncorrelated, satisfy $\text{Cov}(Z_t, Z_{t-1}) = 0$. Recall that $\text{Cov}(Z_t, Z_{t-1}) = E[Z_t Z_{t-1}] - E[Z_t]E[Z_{t-1}]$ and that being uncorrelated is sufficient for the expectation

of the product of two random variables to be equal to the product of their expectations. One can easily verify that for all t

$$\text{Cov}(X_t, X_{t+h}) = \begin{cases} (1 + \theta^2)\sigma^2 & \text{if } h = 0 \\ \theta\sigma^2 & \text{if } h = \pm 1 \\ 0 & \text{if } |h| \geq 2. \end{cases} \quad (2.2)$$

Thus the conditions of weak stationarity hold and the sequence $\{X_t\}$ is weakly stationary. The autocorrelation function is given by:

$$\rho(h) = \frac{\text{Cov}(X_0, X_h)}{\text{Var}(X_0)} = \begin{cases} 1 & \text{if } h = 0 \\ \frac{\theta}{1+\theta^2} & \text{if } h = \pm 1 \\ 0 & \text{if } |h| \geq 2. \end{cases} \quad (2.3)$$

Example 2.5 (First-Order Autoregression). Let $\{X_t\}$ be a stationary series satisfying the equation:

$$X_t = \phi X_{t-1} + Z_t \quad t = 0, \pm 1, \pm 2, \dots, \quad (2.4)$$

where $\{Z_t\} \sim \text{WN}(0, \sigma^2)$, $|\phi| < 1$ and Z_t is uncorrelated with X_s for all $s < t$. As can be seen, $E[X_t] = 0$ and for $h > 0$:

$$\begin{aligned} \text{Cov}(X_{t+h}, X_t) &= \text{Cov}(\phi X_{t+h-1} + Z_{t+h}, X_t) \\ &= \phi \text{Cov}(X_{t+h-1}, X_t) + \text{Cov}(Z_{t+h}, X_t) \\ &= \phi \text{Cov}(X_{t+h-1}, X_t). \end{aligned}$$

From this, one can show that $\text{Cov}(X_{t+h}, X_t) = \phi^h \text{Cov}(X_t, X_t)$. By assumption, $\{X_t\}$ is stationary and hence $\text{Cov}(X_{t+h}, X_t) = \phi^h \text{Cov}(X_0, X_0)$. Next suppose that $h < 0$. Let $s = t + h = t - |h|$ so that $t = s + |h|$. Then $\text{Cov}(X_{t+h}, X_t) = \text{Cov}(X_s, X_{s+|h|}) = \phi^{|h|} \text{Cov}(X_s, X_s) = \phi^{|h|} \text{Cov}(X_0, X_0)$. The autocorrelation function at lag h is

$$\rho(h) = \frac{\phi^{|h|} \text{Cov}(X_0, X_0)}{\text{Cov}(X_0, X_0)} = \phi^{|h|}.$$

One can also obtain a closed-form expression for $\text{Cov}(X_0, X_0)$. Since $\text{Cov}(X_0, X_0) = \text{Cov}(X_t, X_t)$ and

$$\text{Cov}(X_t, X_t) = \text{Cov}(\phi X_{t-1} + Z_t, \phi X_{t-1} + Z_t) = \phi^2 \text{Cov}(X_0, X_0) + \text{Cov}(Z_t, Z_t),$$

we obtain that $\text{Cov}(X_0, X_0) = \sigma^2 / (1 - \phi^2)$.

Chapter 3

Trees, Boosting and Random Forests

Chapter 4

Gradient Boosted Regression

We first consider gradient boosting for classification problems. Suppose that we are given a dataset $\{(\mathbf{x}_i, y_i)_{i=1}^n\}$, where $\mathbf{x}_i \in \mathbf{R}^p$ and $y_i \in \{0, 1\}$. The problem is to find a mapping $f: \mathbf{x} \rightarrow y$. As in logistic regression, we modify the problem slightly and do not work directly with the class labels y_i . Instead, we consider the $\log(\text{odds})$ of the event $P\{y = 1 \mid \mathbf{x}\}$. Since $-\infty < \log(\text{odds}) < +\infty$, this restatement allows us to focus on functions $F: \mathbf{R}^p \rightarrow \mathbf{R}$ rather than from $\mathbf{R}^p \rightarrow \{0, 1\}$. This is helpful because gradient boosting requires differentiable loss functions.

Our next step is to build an appropriate loss function. Let $p = P\{y = 1 \mid \mathbf{x}\}$. Then we may write

$$P\{y \mid \mathbf{x}\} = p^y \cdot (1 - p)^{1-y}. \quad (4.1)$$

Consequently, the likelihood of y_1, \dots, y_n given $\mathbf{x}_1, \dots, \mathbf{x}_n$ assuming that the data $\{(\mathbf{x}_i, y_i)_{i=1}^n\}$ are independent is

$$P\{y_1, \dots, y_n \mid \mathbf{x}_1, \dots, \mathbf{x}_n\} = \prod_{i=1}^n p_i^{y_i} \cdot (1 - p_i)^{1-y_i}. \quad (4.2)$$

The log-likelihood is $\sum_i [y_i \log p_i + (1 - y_i) \log(1 - p_i)]$. If we were to fit a model such a logistic regression, we would search for those model parameters for which the log-likelihood is a maximum. If we were to use a transform of the log-likelihood as a loss function, we would want to minimize that transform. The easiest transform is the negative log-likelihood. Minimizing the negative log-likelihood is equivalent to maximizing the log-likelihood (which is what we want to do). This negative log-likelihood will then be our loss function and this also goes by the name of *cross entropy*.

We re-write the loss function in terms of $\log(\text{odds})$ because this is what our gradient boosting model will output. To simplify the resulting expressions, we consider only one data point and omit the data index i . We may then write:

$$\begin{aligned} -[y \log p + (1 - y) \log(1 - p)] &= -y \log p - (1 - y) \log(1 - p) \\ &= -y \log p + y \log(1 - p) - \log(1 - p) \\ &= -y \log \frac{p}{1 - p} - \log \left(1 - \sigma \left(\log \frac{p}{1 - p} \right) \right). \end{aligned}$$

In the last step, we used the fact that $p = \sigma(p/(1 - p))$, where σ is the sigmoid function. This last step can be simplified by expanding out the sigmoid function and the loss function can then be written as:

$$-y \log \frac{p}{1 - p} + \log(1 + e^{\log \frac{p}{1 - p}}). \quad (4.3)$$

Note the positive sign before the logarithm.

Finally, we note that the gradient boosting engine gives us a function $F(\mathbf{x})$ that represents the $\log(\text{odds})$. With this, we can write the loss function more clearly as:

$$L(y, F(\mathbf{x})) = -yF(\mathbf{x}) + \log(1 + e^{F(\mathbf{x})}). \quad (4.4)$$

This function is differentiable wrt $F(\mathbf{x})$ and with some manipulation, one can show that:

$$\frac{dL}{dF(\mathbf{x})} = -y + \sigma(F(\mathbf{x})). \quad (4.5)$$

Now this has a nice interpretation. The term $\sigma(F(\mathbf{x}))$ is the predicted probability that $y = 1$. If we were to interpret the label $y \in \{0, 1\}$ as a probability, then the derivative of the loss function wrt $F(\mathbf{x})$ is the negative of the difference of the actual probability and the predicted probability.

We now consider the gradient boosting algorithm and go through it step by step.

Input. A dataset $\{(\mathbf{x}_i, y_i)_{i=1}^n\}$ and a differentiable loss function $L(y, F(\mathbf{x}))$.

Step 1. Initialize model with a constant value $F_0(\mathbf{x})$ such that for all \mathbf{x}_j , $1 \leq j \leq n$,

$$F_0(\mathbf{x}_j) = \operatorname{argmin}_{\gamma} \sum_{i=1}^n L(y_i, \gamma). \quad (4.6)$$

Step 2 For $m = 1$ to M (the maximum number of trees):

1. For each $i \in [1, \dots, n]$, compute the pseudo residuals:

$$r_{im} = - \left[\frac{\partial L(y, F(\mathbf{x}))}{\partial F(\mathbf{x})} \right]_{F(\mathbf{x})=F_{m-1}(\mathbf{x}), (\mathbf{x}_i, y_i)}. \quad (4.7)$$

2. For each $i \in [1, \dots, n]$, fit a regression tree to the r_{im} values and create terminal regions R_{jm} , where $j = 1, \dots, J_m$, where J_m is the number of leaves in the m th tree.
3. For each $j \in [1, \dots, J_m]$, compute an output value for leaf j in tree m :

$$\gamma_{jm} = \operatorname{argmin}_{\gamma} \sum_{\mathbf{x}_i \in R_{jm}} L(y_i, F_{m-1}(\mathbf{x}_i) + \gamma). \quad (4.8)$$

4. Update $F_m(\mathbf{x}) = F_{m-1}(\mathbf{x}) + \nu \cdot \sum_{j=1}^{J_m} \gamma_{jm} I(\mathbf{x} \in R_{jm})$.

Step 3 Return F_M .

Step 1 asks us to initialize the model with a constant value γ which is the solution to $\operatorname{argmin}_{\gamma} \sum_{i=1}^n L(y_i, \gamma)$. Now,

$$\sum_{i=1}^n L(y_i, \gamma) = \sum_{i=1}^n (-y_i \gamma + \log(1 + e^{\gamma})) = -\gamma \sum_{i=1}^n y_i + n \cdot \log(1 + e^{\gamma}) \stackrel{\text{def}}{=} g(\gamma). \quad (4.9)$$

Differentiating this wrt γ , we obtain:

$$\frac{dg}{d\gamma} = -\sum_{i=1}^n y_i + n \cdot \frac{e^{\gamma}}{1 + e^{\gamma}} = -\sum_{i=1}^n y_i + n \cdot \sigma(\gamma). \quad (4.10)$$

Setting the right hand expression above to 0, we obtain:

$$\gamma = \log \frac{\bar{y}_n}{1 - \bar{y}_n}. \quad (4.11)$$

Thus the initial constant solution is the log(odds) of the mean $P\{y = 1 \mid \mathbf{x}\}$ in the data. Intuitively, this seems like a good initial solution to start out with.

Step 2 is where all the trees are constructed. There are M trees in total and this number has to be decided beforehand. In practice, this is a parameter to algorithm and is determined using techniques such as cross-validation. The first step in this sequence asks us to compute “residuals” r_{im} for each data point (\mathbf{x}_i, y_i) and each tree. Computing the residuals amounts to computing:

$$r_{im} = - \left[\frac{\partial L(y, F(\mathbf{x}))}{\partial F(\mathbf{x})} \right]_{F(\mathbf{x})=F_{m-1}(\mathbf{x}), (\mathbf{x}_i, y_i)} = y_i - \sigma(F_{m-1}(\mathbf{x}_i)) \quad (4.12)$$

Note that $F_{m-1}(\mathbf{x})$ gives the log(odds) of the event that $y = 1$ given the data \mathbf{x} . Hence $\sigma(F_{m-1}(\mathbf{x}))$ represents $P\{y = 1 \mid \mathbf{x}\}$. Consequently, r_{im} is the difference between the observed probability and the predicted probability. This looks very much like the residuals as defined in linear regression. Thus the name “pseudo residuals.”

In the next step in the sequence for Step 2 is the construction of a regression tree to fit these pseudo residual values. Suppose that the m th tree has J_m leaves. The third step determines an appropriate output value of each leaf of the tree just constructed. The output value for the j th leaf of this tree is:

$$\gamma_{jm} = \operatorname{argmin}_{\gamma} \sum_{\mathbf{x}_i \in R_{jm}} L(y_i, F_{m-1}(\mathbf{x}_i) + \gamma). \quad (4.13)$$

We could differentiate the above expression wrt γ but this is potentially very messy. Instead, we simplify the loss function by using a second-order Taylor expansion. In this context, recall that if $f: \mathbf{R} \rightarrow \mathbf{R}$ is a function that is infinitely differentiable at a point x , then

$$f(x+h) \approx f(x) + \frac{f^{(1)}(x)}{1!} \cdot h + \frac{f^{(2)}(x)}{2!} \cdot h^2 + \frac{f^{(3)}(x)}{3!} \cdot h^3 + \dots \quad (4.14)$$

Using this, we can write a second-order approximation to our loss function:

$$\begin{aligned} L(y_i, F_{m-1}(\mathbf{x}_i) + \gamma) &\approx L(y_i, F_{m-1}(\mathbf{x}_i)) + \gamma \cdot \frac{dL(y_i, F_{m-1}(\mathbf{x}_i))}{dF_{m-1}} + \frac{\gamma^2}{2} \cdot \frac{d^2 L(y_i, F_{m-1}(\mathbf{x}_i))}{dF_{m-1}^2} \\ &= L(y_i, F_{m-1}(\mathbf{x}_i)) + \gamma \cdot (-y_i + \sigma(F_{m-1}(\mathbf{x}_i))) + \frac{\gamma^2}{2} \cdot \sigma(F_{m-1}(\mathbf{x}_i))(1 - \sigma(F_{m-1}(\mathbf{x}_i))). \end{aligned}$$

Differentiate the right-hand side wrt γ to obtain:

$$-y_i + \sigma(F_{m-1}(\mathbf{x}_i)) + \gamma \cdot \sigma(F_{m-1}(\mathbf{x}_i))(1 - \sigma(F_{m-1}(\mathbf{x}_i))). \quad (4.15)$$

Set this to 0 and solve for γ :

$$\gamma = \frac{y_i - \sigma(F_{m-1}(\mathbf{x}_i))}{\sigma(F_{m-1}(\mathbf{x}_i))(1 - \sigma(F_{m-1}(\mathbf{x}_i)))} = \frac{y_i - p_i}{p_i(1 - p_i)}. \quad (4.16)$$

This is the expression for just a single data point. Taking into account all data points, we obtain:

$$\gamma = \frac{\sum_{i=1}^n (y_i - \sigma(F_{m-1}(\mathbf{x}_i)))}{\sum_{i=1}^n \sigma(F_{m-1}(\mathbf{x}_i))(1 - \sigma(F_{m-1}(\mathbf{x}_i)))} = \frac{\sum_i (y_i - p_i)}{\sum_i p_i(1 - p_i)}. \quad (4.17)$$

We can now evaluate the value γ_{jm} of each leaf node of the m th tree. The fourth and final step of Step 2 is to update the prediction function F_{m-1} using the values of the leaf nodes of the tree just constructed.

$$F_m(\mathbf{x}) = F_{m-1}(\mathbf{x}) + \nu \cdot \sum_{j=1}^{J_m} \gamma_{jm} I(\mathbf{x} \in R_{jm}).$$

Step 3 simply outputs the function F_M obtained after the updatation in the M th round.

Chapter 5

Neural Networks

Bibliography

- [1] Joseph K. Blitzstein, Jessica Hwang. *Introduction to Probability*, Second Edition, Chapman and Hall, 2019.
- [2] Peter J. Brockwell, Richard A. Davis. *Introduction to Time Series and Forecasting*, Third Edition, Springer, 2016.
- [3] George Casella, Roger L. Berger. *Statistical Inference*, Second Edition, Duxbury Advanced Series, 2001.
- [4] Morris DeGroot, Mark J. Schervish. *Probability and Statistics*, Fourth Edition, Pearson, 2012.
- [5] Rob J. Hyndman and George Athanasopoulos. *Forecasting: Principles and Practice*, Second Edition. Online book at: <https://otexts.com/fpp2/>
- [6] Ben Lambert. *A Student's Guide to Bayesian Statistics*, Sage Publications, 2018.