

Statistics 2 - Notes

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1 Estimation

1.1 Introduction

Definition 1.1 - *Probabilty Space, $(\Omega, \mathcal{F}, \mathbb{P})$*

A mathematical construct for modelling the real world. A *Probabilty Space* has three elements

- i) Ω - Sample space.
- ii) \mathcal{F} - Set of events.
- iii) \mathbb{P} - Probability measure.

and most fulfil the following conditions

- i) $\Omega \in \mathcal{F}$;
- ii) $\forall A \in \mathcal{F} \implies A^c \in \mathcal{F}$;
- iii) $\forall A_0, \dots, A_n \in \mathcal{F} \implies \left(\bigcup_i A_i \right) \in \mathcal{F}$;
- iv) $\mathbb{P}(\Omega) = 1$; and,
- v) $\mathbb{P}\left(\bigcup_{i=1}^{\infty} A_i\right) = \sum_{i=1}^{\infty} \mathbb{P}(A_i)$ for disjoint A_1, A_2, \dots (Countable Additivity).

Definition 1.2 - *Random Variable*

A function which maps an event in the sample space to a value *e.g.* $X : \Omega \rightarrow \mathbb{R}$.

Remark 1.1 - *Probability Density Function for iid Random Variable Vector*

For $\mathbf{X} \sim f_n(\cdot; \theta)$ where each component of \mathbf{X} is independent and identically distribution the probability density function of \mathbf{X} is

$$f_n(\mathbf{x}; \theta) = \prod_{i=1}^n f(x_i; \theta)$$

Definition 1.3 - *Expectation*

The mean value for a random variable. For rv X

$$\mathbb{E}(X) := \sum_{x \in \mathcal{X}} x f_X(x) \quad \& \quad \mathbb{E}(X) := \int_{\mathbb{R}} x f_X(x) dx$$

Theorem 1.1 - *Expection of a Function*

For a function $g : \mathbb{R} \rightarrow \mathbb{R}$ and rv X with pmf f_X

$$\mathbb{E}(g(X)) := \sum_{g(x) \in \mathcal{X}} x f_X(x) \quad \& \quad \mathbb{E}(g(X)) := \int_{\mathbb{R}} g(x) f_X(x) dx$$

Theorem 1.2 - *Expectation of a Linear Operator*

For rv X with pmf f_X & $a, b \in \mathbb{R}$

$$\mathbb{E}(aX + b) = a\mathbb{E}(X) + b$$

Definition 1.4 - *Variance*

For rv X

$$\text{Var}(X) := \mathbb{E}[(X - \mathbb{E}(X))^2] = \mathbb{E}(X^2) - \mathbb{E}(X)^2$$

Theorem 1.3 - Variance of a Linear Operator

For rv X and $a, b \in \mathbb{R}$

$$\text{Var}(aX + b) = a^2 \text{Var}(X)$$

Definition 1.5 - Moment of a Random Variable

For rv X the n^{th} moment of X is defined as $\mathbb{E}(X^n)$.

N.B. - $\mathbb{E}(X^n) \neq \mathbb{E}(X)^n$.

Definition 1.6 - Covariance

For rv X & Y

$$\text{cov}(X, Y) := \mathbb{E}[(X - \mathbb{E}(X))(Y - \mathbb{E}(Y))] = \mathbb{E}(XY) - \mathbb{E}(X)\mathbb{E}(Y)$$

Theorem 1.4 - Properties of Covariance

Let X & Y be independent random variables

$$\text{i) } \text{cov}(X, X) = \text{Var}(X);$$

$$\text{ii) } \text{cov}(X, Y) = 0$$

Theorem 1.5 - Variance of two Random Variables with linear operators

$$\text{Var}(aX + bY) = a^2 \text{Var}(X) + b^2 \text{Var}(Y) + 2ab \text{cov}(X, Y)$$

Theorem 1.6 - Independent Random Variables

Random variables X_1, \dots, X_n are independent iff

$$\mathbb{P}(X_1 \leq a_1, \dots, X_n \leq a_n) = \prod_{i=1}^n \mathbb{P}(X_i \leq a_i) \quad \forall a_1, \dots, a_n \in \mathbb{R}$$

2 The Likelihood Function

Definition 2.1 - Likelihood Function

Define $\mathbf{X} \sim f_n(\cdot; \theta^*)$ for some unknown $\theta^* \in \Theta$ and let \mathbf{x} be an observation of \mathbf{X} .

A *Likelihood Function* is any function, $L(\cdot; \mathbf{x}) : \Theta \rightarrow [0, \infty)$, which is proportional to the PMF/PDF of the observed realisation \mathbf{x} .

$$L(\theta; \mathbf{x}) := C f_b(\mathbf{x}; \theta) \quad \forall C > 0$$

N.B. Sometimes this is called the *Observed Likelihood Function* since it is dependent on observed data.

Definition 2.2 - Log-Likelihood Function

Let $\mathbf{X} \sim f_n(\cdot; \theta^*)$ for some unknown $\theta^* \in \Theta$ and \mathbf{x} be an observation of \mathbf{X} .

The *Log-Likelihood Function* is the natural log of a *Likelihood Function*

$$\ell(\theta; \mathbf{x}) := \ln f_n(\mathbf{x}; \theta) + C, \quad C \in \mathbb{R}$$

Theorem 2.1 - Multidimensional Transforms

Let \mathbf{X} be a continuous random vector in \mathbb{R}^n with PDF $f_{\mathbf{X}}$; $g : \mathbb{R}^n \rightarrow \mathbb{R}^n$ be a continuous differentiable bijection; and, $h := g^{-1}$.

Then $\mathbf{Y} = g(\mathbf{X})$ is a continuous random vector and its PDF is

$$f_{\mathbf{Y}}(\mathbf{y}) = f_{\mathbf{X}}(h(\mathbf{y}))H_h(\mathbf{Y})$$

where

$$J_h := \left| \det \left(\frac{\partial h}{\partial \mathbf{y}} \right) \right|$$

Proposition 2.1 - *Invariance of Likelihood Function by bijective transformation of the observations independent of θ*

Let $g : \mathbb{R}^n \rightarrow \mathbb{R}^n$ be a bijective transformation which is independent of θ ; and $\mathbf{Y} := g(\mathbf{X})$.

Then \mathbf{Y} is a random variable with PDF/PMF

$$f_{\mathbf{Y}}(\mathbf{y}; \theta) \propto f_{\mathbf{X}}(g^{-1}(\mathbf{y}); \theta)$$

Hence, if $\mathbf{y} = g(\mathbf{x})$ then $L_{\mathbf{Y}}(\theta; \mathbf{y}) \propto L_{\mathbf{X}}(\theta; \mathbf{x})$

Proof 2.1 - *Proposition 2.1*

Let $g : \mathbb{R}^n \rightarrow \mathbb{R}^n$ be a bijective transformation which is independent of θ ; $h := g^{-1}$; \mathbf{X}, \mathbf{Y} be a rvs st $\mathbf{Y} := g(\mathbf{X})$.

i) *Discrete Case* - Consider the case when \mathbf{X} is a discrete rv. Then

$$\begin{aligned} f_{\mathbf{Y}}(\mathbf{y}; \theta) &= \mathbb{P}(\mathbf{Y} = \mathbf{y}; \theta) \\ &= \mathbb{P}(g^{-1}(\mathbf{Y}) = g^{-1}(\mathbf{y}); \theta) \\ &= \mathbb{P}(h(\mathbf{Y}) = h(\mathbf{y}); \theta) \\ &= \mathbb{P}(\mathbf{X} = h(\mathbf{y}); \theta) \\ &= f_{\mathbf{X}}(g^{-1}(\mathbf{y}); \theta) \end{aligned}$$

ii) *Continuous Case* - Consider the case when \mathbf{X} is a continuous rv. Then, by **Theorem 2.1**

$$f_{\mathbf{Y}}(\mathbf{y}; \theta) = f_{\mathbf{X}}(g^{-1}(\mathbf{y}); \theta) J_{g^{-1}}(\mathbf{y})$$

Since $J_{g^{-1}}$ does not depend on θ this case is solved.

Thus in both cases $L_{\mathbf{Y}}(\theta; \mathbf{y}) = f_{\mathbf{Y}}(\mathbf{y}; \theta) \propto f_{\mathbf{X}}(g^{-1}(\mathbf{y}); \theta) = L_{\mathbf{X}}(\theta; \mathbf{x})$.

3 Maximum Likelihood Estimates

Definition 3.1 - *Maximum Likelihood Estimate*

Let $\mathbf{X} \sim f_n(\cdot; \theta)$; and \mathbf{x} be a realisation of \mathbf{X} .

The *Maximum Likelihood Estimate* is the value $\hat{\theta} \in \Theta$ st

$$\forall \theta \in \Theta \quad f_n(\mathbf{x}; \hat{\theta}) \geq f_n(\mathbf{x}, \theta)$$

Equivalently

$$\forall \theta \in \Theta \quad L(\hat{\theta}; \mathbf{x}) \geq L(\theta; \mathbf{x}) \quad \text{or} \quad \ell(\hat{\theta}; \mathbf{x}) \geq \ell(\theta; \mathbf{x})$$

i.e. $\hat{\theta}(\mathbf{x}) := \arg\max_{\theta} (L(\theta; \mathbf{x}))$.

Remark 3.1 - *The Maximum Likelihood Estimate may not be unique*

Example 3.1 - *MLE for Uniform Distribution*

Consider $\mathbf{X} \stackrel{\text{iid}}{\sim} U[0, \theta]$ for $\theta > 0$.

Then

$$\begin{aligned} L(\theta; \mathbf{x}) &\propto f_n(\mathbf{x}; \theta) \\ &= \prod_{i=1}^n \frac{1}{\theta} \mathbb{1}\{x_i \in [0, \theta]\} \\ &= \frac{1}{\theta^n} \prod_{i=1}^n \mathbb{1}\{x_i \in [0, \theta]\} \\ \implies \hat{\theta} &= \max\{x_i : x_i \in \mathbf{x}\} \end{aligned}$$

Remark 3.2 - MLE of Reparameterisation

Define $\tau(\theta) : \mathbb{R} \rightarrow \mathbb{R}$. Then

$$\hat{\tau} = \tau(\hat{\theta})$$

N.B. We often write \tilde{f} to represent the pmf when τ is taken as a parameter rather than θ . i.e. $f(x; \theta) = \tilde{f}(x; \tau(\theta))$.

Theorem 3.1 - Invariance of MLE under bijective Reparameterisation

Let $g : \Theta \rightarrow G$ be a bijective transformation of the statistical parameter θ .

Let $\mathbf{X} \sim f(\cdot; \theta) = \tilde{f}(\cdot; g(\theta))$ for some θ , and let \mathbf{x} be a realisation of \mathbf{X} .

If $\hat{\theta}$ is an MLE of θ then $\hat{\tau} = g(\hat{\theta})$ is an MLE of τ .

Proof 3.1 - Theorem 3.1

This is a proof by contradiction.

Suppose $\exists \tau^* \in G$ s.t. $\tilde{f}(x; \tau^*) > \tilde{f}(x; \hat{\tau})$. We know that $\forall \theta \in \Theta$, $f(x; \theta) = \tilde{f}(x; g(\theta))$ and $\forall \tau \in G$, $f(x; g^{-1}(\tau)) = \tilde{f}(x; \tau)$.

We deduce that

$$\begin{aligned} f(x; g^{-1}(\tau^*)) &= \tilde{f}(x; \tau^*) \\ &> \tilde{f}(x; \hat{\tau}) \text{ by assumption} \\ &= f(x; g^{-1}(\hat{\tau})) \\ &= f(x; \hat{\theta}) \end{aligned}$$

This contradicts the assumption that $\hat{\theta}$ is a maximum likelihood estimate of θ .

Remark 3.3 - Not all Reparameterisations are Bijective

When reparameterisations $g : \mathbb{R} \rightarrow \mathbb{R}$ is not bijective it is helpful to consider the *induced likelihood*

$$L^*(\tau; \mathbf{x}) := \max_{\theta \in G_\tau} L(\theta; \mathbf{x}) \text{ where } G_\tau := \{\theta : g(\theta) = \tau\}$$

Since this reduces the domain to only where g is bijective.

3.1 Determining MLEs - The Tractable Case**Proposition 3.1 - Differentiable Likelihood in the continuous case - Multivariate**

When $L(\theta; \mathbf{x})$ is differentiable one can find MLEs by considering its extrema. This is done equating & solving the cases when the gradient is zero, i.e. $\nabla L(\theta; \mathbf{x}) = 0$, and then checking whether this is a maximum or minimum point.

A point is a local minimum if the Hessian at the point is *Negative Definite* i.e. $x^T A x < 0 \forall x \neq \mathbf{0}$.

Example 3.2 - MLE of Normal Distribution

Let $\mathbf{X} \stackrel{\text{iid}}{\sim} \mathcal{N}(\mu, \sigma^2)$

$$\begin{aligned} L(\mu, \sigma^2; \mathbf{x}) &= \prod_{i=1}^n \frac{1}{\sqrt{2\pi\sigma^2}} e^{-\frac{(x_i - \mu)^2}{2\sigma^2}} \\ \Rightarrow \ell(\mu, \sigma^2; \mathbf{x}) &= C - \frac{n}{2} \ln(\sigma^2) - \frac{1}{2\sigma^2} \sum_{i=1}^n (x_i - \mu)^2 \\ \Rightarrow \nabla \ell(\mu, \sigma^2; \mathbf{x}) &= \left(\frac{-1}{\sigma^2} \sum_{i=1}^n (x_i - \mu), \quad -\frac{n}{2\sigma^2} + \frac{1}{2\sigma^4} \sum_{i=1}^n (x_i - \mu)^2 \right) \\ \text{Setting } \frac{-1}{\sigma^2} \sum_{i=1}^n (x_i - \mu) &= 0 \\ \Rightarrow \hat{\mu} &= \frac{1}{n} \sum_{i=1}^n x_i = \bar{x} \\ \text{Setting } -\frac{n}{2\sigma^2} + \frac{1}{2\sigma^4} \sum_{i=1}^n (x_i - \mu)^2 &= 0 \\ \Rightarrow \hat{\sigma}^2 &= \frac{1}{n} \sum_{i=1}^n (x_i - \hat{\mu})^2 \end{aligned}$$

We now want to check whether $(\hat{\mu}, \hat{\sigma}^2)$ is a minimum.

$$\begin{aligned}\nabla^2 \ell(\mu, \sigma^2; \mathbf{x}) &= \begin{pmatrix} \frac{\partial^2 \ell(\mu, \sigma^2; \mathbf{x})}{\partial \mu^2} & \frac{\partial^2 \ell(\mu, \sigma^2; \mathbf{x})}{\partial \mu \partial \sigma^2} \\ \frac{\partial^2 \ell(\mu, \sigma^2; \mathbf{x})}{\partial \mu \partial \sigma^2} & \frac{\partial^2 \ell(\mu, \sigma^2; \mathbf{x})}{\partial (\sigma^2)^2} \end{pmatrix} \\ &= \begin{pmatrix} -\frac{n}{\hat{\sigma}^2} & 0 \\ 0 & -\frac{n}{2\hat{\sigma}^4} \end{pmatrix}\end{aligned}$$

Since $\begin{pmatrix} z_1 & z_2 \end{pmatrix} \begin{pmatrix} -a & 0 \\ 0 & -b \end{pmatrix} \begin{pmatrix} z_1 \\ z_2 \end{pmatrix} = -az_1^2 - bz_2^2 < 0 \forall a, b > 0$ and we have $\frac{n}{\hat{\sigma}^2}, \frac{n}{2\hat{\sigma}^4} > 0$ then we can conclude that $\nabla^2 \ell$ is negative definite.

Thus $\hat{\mu} = \bar{x}$ & $\hat{\sigma}^2 = \frac{1}{n} \sum_{i=1}^n (x_i - \hat{\mu})^2$ is an MLE for the normal distribution.

Example 3.3 - MLE for Capture-Recapture Model

Suppose you are wanting to calculate the unknown size of a population, n . The Capture-Recapture Model is one technique that can be used. You tag $t \leq n$ members of the population; wait for a while; then recapture $c \leq n$ members of which $x \leq \min\{t, c\} \leq n$ are tagged.

With t, c, x known produce a MLE for n .

We first work out the associated probability distribution for X , the population size. We have

- i) $\binom{t}{x}$ ways of choosing x members among the tagged ones;
- ii) $\binom{n-t}{c-x}$ ways of choosing the remaining members among the non-tagged ones;
- iii) $\binom{n}{c}$ ways of choosing c members in a population of n individuals.

Thus

$$f_X(x; n) = \frac{\binom{t}{x} \binom{n-t}{c-x}}{\binom{n}{c}}$$

This means that $X \sim \text{Hypergeometric}(t, n, c)$ with t & c known.

Now we calculate the MLE for X

$$\begin{aligned}L(n; x) &= f_X(x; n) \\ &= \frac{\binom{t}{x} \binom{n-t}{c-x}}{\binom{n}{c}} \\ &= \frac{t!}{x!(t-x)!} \frac{(n-t)!}{(c-x)!(n-t-c+x)!} \\ &= \frac{n!}{c!(n-c)!}\end{aligned}$$

Now we consider $L(n; x) = 0$ when $x > \min\{t, c\}$. We want to identify values of n for which $L(n; x) \geq L(n-1; x)$.

Consider $n - 1 \geq \min\{t, c\} \implies L(n - 1; x) > 0$

$$\begin{aligned}
 \text{Let } r(n) &:= \frac{L(n; x)}{L(n - 1; x)} \\
 &= \frac{n - t}{n - t - c + x} \frac{n - c}{n} \\
 \implies 1 &\leq r(n) \\
 \Leftrightarrow 1 &\leq \frac{n - t}{n - t - c + x} \frac{n - c}{n} \\
 \Leftrightarrow n(n - t - c + x) &\leq (n - t)(n - c) \\
 \Leftrightarrow n^2 - nt - cn + xn &\leq n^2 - nt - cn + ct \\
 \Leftrightarrow xn &\leq ct \\
 \Leftrightarrow x &\leq \frac{ct}{n}
 \end{aligned}$$

So $L(n; x)$ is increasing for $n \leq \lfloor \frac{ct}{x} \rfloor$ & decreasing for $n > \lfloor \frac{ct}{x} \rfloor$.
 Consequently $\hat{n}_{\text{MLE}}(x) = \lfloor \frac{ct}{x} \rfloor$

4 Statistics and Estimators

Definition 4.1 - Statistic

Given some data \mathbf{x} a statistic is a function of the data $T(\mathbf{x})$.

N.B. A statistic cannot depend on an unknown statistical parameter.

Definition 4.2 - Estimate

Let $\mathbf{X} \sim f_n(\cdot; \theta^*)$ with $\theta^* \in \Theta$ and \mathbf{x} be a realisation of \mathbf{X} .

An *Estimate* θ^* is a statistic $\hat{\theta}(\mathbf{x}) = T(\mathbf{x})$ which is intended to approximate the real value of θ^* .

N.B. An *Estimate* is a real value & thus is hard to evaluate.

Definition 4.3 - Estimator

Let $\mathbf{X} \sim f_n(\cdot; \theta^*)$ with $\theta^* \in \Theta$ and \mathbf{x} be a realisation of \mathbf{X} .

An *Estimator* of θ^* is $\hat{\theta}$ where $\hat{\theta}(\mathbf{x})$ is an *estimate*.

N.B. We call $T(\mathbf{X})$ an estimator. This is a random variable.

Definition 4.4 - Distribution of an Estimator

Let $\mathbf{X} | \text{sim} f_n(\cdot; \theta^*)$ with $\theta^* \in \Theta \subseteq \mathbb{R}$.

If $\hat{\theta}(\mathbf{X})$ is a real-valued random variable, we can write its CDF as

$$\begin{aligned}
 F_{\hat{\theta}(\mathbf{X})}(t; \theta^*) &= \mathbb{P}(\hat{\theta}(\mathbf{X}) \leq t; \theta^*) \\
 &= \int_{\mathcal{X}^n} \mathbb{1}\{\hat{\theta}(\mathbf{x}) \leq t\} f_n(\mathbf{x}; \theta^*) d\mathbf{x}
 \end{aligned}$$

Remark 4.1 - Estimator depends upon true value

The distribution of $\hat{\theta}(\mathbf{X})$ depends on the distribution of \mathbf{X} which in turn depends upon the distribution of θ^* .

Thus the distribution of an estimator depends on the true parameter of the variable it is estimating.

Remark 4.2 - Estimator Distribution & Sample Size

As sample size increases the distribution of an estimator may converge to a more standard distribution (e.g. Normal, Poisson).

Definition 4.5 - Bias

Bias is a measure of how much an estimator deviates from the true value, on average.

$$\begin{aligned}\text{Bias}(\hat{\theta}; \theta^*) &:= \mathbb{E}(\hat{\theta}(\mathbf{X}) - \theta^*; \theta^*) \\ &= \mathbb{E}(\hat{\theta}; \theta^*) - \mathbb{E}(\theta^*; \theta^*) \\ &= \mathbb{E}(\hat{\theta}; \theta^*) - \theta^*\end{aligned}$$

Definition 4.6 - Unbiased Estimator

An *Estimator*, $\hat{\theta}$, is said to be *Unbiased* if $\forall \theta \in \Theta$, $\text{Bias}(\hat{\theta}; \theta) = 0$.
Equivalently $\mathbb{E}(\hat{\theta}; \theta) = \theta$.

Definition 4.7 - Mean Square Error

The *Mean Square Error* of an estimator is the mean of the squared error associated with rv $\hat{\theta}$.

$$MSE(\hat{\theta}; \theta^*) := \mathbb{E} \left[(\hat{\theta}(\mathbf{X}) - \theta^*)^2; \theta^2 \right]$$

Proposition 4.1 - Simplification of MSE Formula

The MSE is a combination of variance & bias.

$$\begin{aligned}MSE(\hat{\theta}; \theta^*) &= \mathbb{E} \left[(\hat{\theta}(\mathbf{X}) - \theta^*)^2; \theta^2 \right] \\ &= \mathbb{E} \left[\left\{ \hat{\theta} - \mathbb{E}(\hat{\theta}; \theta^*) \right\}^2; \theta^* \right] + \left(\mathbb{E}(\hat{\theta} - \theta^*; \theta^*) \right)^2 \\ &= \text{Var}(\hat{\theta}; \theta^*) + \text{Bias}(\hat{\theta}; \theta^*)^2\end{aligned}$$

Example 4.1 - Sample mean as an Estimator

Let $\mathbf{X} \stackrel{\text{iid}}{\sim} \text{Poisson}(\lambda^*)$.

Suppose we are using the sample mean, $\hat{\lambda}(\mathbf{x}) := \frac{1}{n} \sum_{i=1}^n x_i$, as an estimate of λ^* . We first want to show this estimator is *Unbiased*

$$\begin{aligned}\mathbb{E}(\hat{\lambda}; \lambda) &= \mathbb{E} \left(\frac{1}{n} \sum_{i=1}^n X_i; \lambda \right) \\ &= d_n^1 \sum_{i=1}^n \mathbb{E}(X_i; \lambda) \\ &= \frac{1}{n} n \lambda \\ &= \lambda\end{aligned}$$

Thus $\hat{\lambda}$ is unbiased.

Now we consider the MSE of $\hat{\lambda}$

$$\begin{aligned}MSE(\hat{\lambda}; \lambda) &= \text{Var}(\hat{\lambda}; \lambda) \\ &= \text{Var} \left(\frac{1}{n} \sum_{i=1}^n X_i; \lambda \right) \\ &= \frac{1}{n^2} \sum_{i=1}^n \text{Var}(X_i; \lambda) \\ &= \frac{1}{n^2} n \lambda \\ &= \frac{\lambda}{n}\end{aligned}$$

This shows that as the sample size increases the MSE of $\hat{\lambda}$ converges to 0.

5 Probabilistic Convergence

Remark 5.1 - Motivation

Here we consider the properties of a maximum likelihood estimators as the sample size increases.

Theorem 5.1 - Markov's Inequality

For a *non-negative* random variable X and a constant $a > 0$

$$\mathbb{P}(X \geq a) \leq \frac{\mathbb{E}(X)}{a}$$

Proof 5.1 - Markov's Inequality

Consider continuous X . We have

$$\begin{aligned} a\mathbb{P}(X \geq a) &= a \int_a^\infty f_X(x) dx \\ &\leq \int_a^\infty x f_X(x) dx \\ &\leq \int_0^\infty x f_X(x) dx \\ &= \mathbb{E}(X) \\ \implies a\mathbb{P}(X \geq a) &= \mathbb{E}(X) \\ \implies \mathbb{P}(X \geq a) &\leq \frac{\mathbb{E}(X)}{a} \end{aligned}$$

□

Theorem 5.2 - Chebyshev's Inequality

Let $\mu = \mathbb{E}(X)$ and $\sigma^2 = \text{Var}(X)$. Then

$$\forall a > 0, \mathbb{P}(|X - \mu| \geq a) \leq \frac{\sigma^2}{a^2}$$

Proof 5.2 - Chebyshev's Inequality

We have

$$\begin{aligned} \mathbb{P}(|X - \mu| \geq a) &= \mathbb{P}(|X - \mu|^2 \geq a^2) \\ &\leq \frac{\mathbb{E}((X - \mu)^2)}{a^2} \text{ By Markov's Inequality} \\ &= \frac{\sigma^2}{a^2} \end{aligned}$$

□

Definition 5.1 - Convergence in Probability

We say the sequence of random variables $\{Z_n\}_{n \in \mathbb{N}}$ converges in probability to the random variable Z if

$$\forall \varepsilon > 0, \lim_{n \rightarrow \infty} \mathbb{P}(|Z_n - Z| > \varepsilon) = 0$$

N.B. This is denoted $Z_n \rightarrow_{\mathbb{P}} Z$.

N.B. The random variables $\{Z_n\}_{n \in \mathbb{N}}$ & Z must be in the same probability space.

Theorem 5.3 - Weak Law of Large Numbers

If $\{X_n\}_{n \in \mathbb{N}}$ are independent & identically distributed and $\mathbb{E}(X_1) = \mu < \infty$ then

$$Z_n = \frac{1}{n} \sum_{i=1}^n X_i \rightarrow_{\mathbb{P}} \mu$$

N.B. This is an example of Convergence in Probability.

Definition 5.2 - Convergence in Distribution

We say the sequence of random variables $\{Z_n\}_{n \in \mathbb{N}}$ converges in distribution to random variable Z if

$$\forall z \in \mathbb{R} \text{ where } \mathbb{P}(Z \leq z) \text{ is continuous, } \lim_{n \rightarrow \infty} \mathbb{P}(Z_n \leq z) = \mathbb{P}(Z \leq z)$$

N.B. This is denoted $Z_n \rightarrow_{\mathcal{D}} Z$.

N.B. The random variables $\{Z_n\}_{n \in \mathbb{N}}$ & Z need not be in the same probability space.

Remark 5.2 - *Equivalent Statements to Convergence in Distribution*

Saying that $Z_n \rightarrow_{\mathcal{D}} Z$ is equivalent to saying that

$$\forall z \in \mathbb{R} \text{ where } F_Z(z) \text{ is continuous, } \lim_{n \rightarrow \infty} F_{Z_n}(z) = F_Z(z)$$

Theorem 5.4 - *Central Limit Theorem*

If $\{X_n\}_{n \in \mathbb{N}}$ are independent & identically distributed, $\mathbb{E}(X_1) = \mu < \infty$ and $\text{Var}(X_1) = \sigma^2 < \infty$ then

$$\frac{\sqrt{n}}{\sigma}(Z_n - \mu) \rightarrow_{\mathcal{D}} Z \sim \text{Normal}(0, 1)$$

Theorem 5.5 - *Convergence in Probability & Distribution*

Convergence in probability \implies Convergence in distribution, **but** the opposite is not necessarily true.

Theorem 5.6 - *Convergence in Probability & Distribution to a Constant*

Convergence in distribution to a constant **and** convergence in probability to a constant are equivalent.

Example 5.1 -

Let $X \sim \text{Bernoulli}(\frac{1}{2})$ and $\{X_n\}_{n \in \mathbb{N}}$ be a sequence of random variables where $X_i := (1 - X) + \frac{1}{n}$. We have

$$F_X(x) = \begin{cases} 0 & , x < 0 \\ \frac{1}{2} & , x \in [0, 1) \\ 1 & , x \geq 1 \end{cases} \quad F_{X_n}(x) = \begin{cases} 0 & , x < \frac{1}{n} \\ \frac{1}{2} & , x \in [\frac{1}{n}, 1 + \frac{1}{n}) \\ 1 & , x \geq 1 + \frac{1}{n} \end{cases}$$

Clearly $F_{X_n}(x) \rightarrow F_X(x)$ at all points at which F_X is continuous (*i.e.* $x \in \mathbb{R} \setminus \{0, 1\}$). Thus $X_n \rightarrow_{\mathcal{D}} X$.

Theorem 5.7 - *Continuous Mapping Theorem*

Let $g : Z \rightarrow G$ be a *continuous* function. Then

- i) If $Z_n \rightarrow_{\mathbb{P}} Z$, then $g(Z_n) \rightarrow_{\mathbb{P}} g(Z)$;
- ii) If $Z_n \rightarrow_{\mathcal{D}} Z$, then $g(Z_n) \rightarrow_{\mathcal{D}} g(Z)$

Theorem 5.8 - *Slutsky's Theorem*

Let $\{Y_n\}_{n \in \mathbb{N}}$ & $\{Z_n\}_{n \in \mathbb{N}}$ be sequences of random variables, Y be a random variable & $c \in \mathbb{R} \setminus 0$ be a constant.

If $Y_n \rightarrow_{\mathcal{D}} Y$ and $Z_n \rightarrow_{\mathcal{D}} c$, then

- i) $Y_n + Z_n \rightarrow_{\mathcal{D}} Y + c$;
- ii) $Y_n Z_n \rightarrow_{\mathcal{D}} Yc$; and,
- iii) $\frac{Y_n}{Z_n} \rightarrow_{\mathcal{D}} \frac{Y}{c}$.

Definition 5.3 - *Convergence in Quadratic Mean*

Let $\{Z_n\}_{n \in \mathbb{N}}$ be a sequence of random variables & Z be a random variable.

We say that $\{Z_n\}_{n \in \mathbb{N}}$ *Converges in Quadratic Mean* to the random variable Z if

$$\lim_{n \rightarrow \infty} \mathbb{E}[(Z_n - Z)^2] = 0$$

N.B. This is denoted $Z_n \rightarrow_{qm} Z$.

Theorem 5.9 - If $Z_n \rightarrow_{qm} Z$ then $Z_n \rightarrow_{\mathbb{P}} Z$

Proof 5.3 - Theorem 5.9

Fix any $\varepsilon > 0$. We have

$$\begin{aligned} \mathbb{P}(|Z_n - Z| > \varepsilon) &= \mathbb{P}(|Z_n - Z|^2 > \varepsilon^2) \\ &\leq \frac{1}{\varepsilon^2} \mathbb{E}[(Z_n - Z)^2] \text{ by Markov's Inequality} \\ &\rightarrow 0 \text{ since } Z_n \rightarrow_{qm} Z. \end{aligned}$$

Hence $Z_n \rightarrow_{\mathbb{P}} Z$.

5.1 Probabilistic Convergence & Estimators

Definition 5.4 - Consistency of a Sequence of Estimators

A sequence of estimators, $\{\hat{\theta}_n(\cdot) : \chi^n \rightarrow \Theta\}$, are said to be *Consistent* if

$$\forall \theta \in \Theta \text{ with } \mathbf{X}_n \sim f_n(\cdot; \theta), \hat{\theta}_n(\mathbf{X}_n) \rightarrow_{\mathbb{P}(\cdot; \theta)} \theta$$

Remark 5.3 - Consistency of a Sequence of Estimators

- i) In numerous situations one will talk about the consistency of *the* estimator, *e.g.* for the MLE, but also for the mean, etc. This implicitly refers to the corresponding sequence of MLEs, sequence of means, etc.
- ii) Note the $\mathbb{P}(\cdot; \theta)$ in the limit above, and in particular the dependence on θ . This is often omitted in practice, you should however not forget what the symbols actually mean.
- iii) Quadratic mean / Mean Square convergence \implies consistency.
That is, if the MSE of the estimator converges to 0, the estimator is consistent.

Example 5.2 - Consistency of Flipping Coins

Let $\mathbf{X}^{\text{iid}} \sim \text{Bernoulli}(\theta^*)$ for some $\theta^* \in [0, 1]$.

The maximum likelihood estimate and method of moments for $\hat{\theta}_n$ are the sample mean.

$$\hat{\theta}_n(X_1, \dots, X_n) = \frac{1}{n} \sum_{i=1}^n X_i$$

By the *Weak Law of Large Numbers* we have that *consistency* of $\{\hat{\theta}_n\}$, since $\mathbb{E}(X_1) = \theta^*$.

Example 5.3 - Crude Confidence Interval when Flipping Coins

Let $\mathbf{X}^{\text{iid}} \sim \text{Bernoulli}(\theta^*)$ for some $\theta^* \in [0, 1]$ and define $\hat{\theta}_n := \hat{\theta}_n(X_1, \dots, X_n)$.

We shall produce a *confidence interval* for θ^* .

$$\begin{aligned} \mathbb{E}(\hat{\theta}_n; \theta^*) &= \theta^* \quad \text{and} \quad \text{Var}(\hat{\theta}_n; \theta^*) = \frac{\theta^*(1 - \theta^*)}{n} \\ \mathbb{P}\left(|\hat{\theta}_n - \theta^*| \geq \varepsilon; \theta^*\right) &\leq \frac{\theta^*(1 - \theta^*)}{n\varepsilon^2} \text{ by Chebyshev's Inequality} \\ \text{We don't know } \theta^*, \text{ but can deduce that } \theta^*(1 - \theta^*) &\leq \frac{1}{4} \\ \implies \mathbb{P}\left(|\hat{\theta}_n - \theta^*| \geq \varepsilon; \theta^*\right) &\leq \frac{1}{4n\varepsilon^2} \\ \text{Define } \alpha &:= \frac{1}{4n\varepsilon^2} \\ \implies \mathbb{P}\left(|\hat{\theta}_n - \theta^*| \geq \frac{1}{2\sqrt{n\alpha}}; \theta^*\right) &\leq \alpha \\ \implies \mathbb{P}\left(\hat{\theta}_n - \frac{1}{2\sqrt{n\alpha}} < \theta^* < \hat{\theta}_n + \frac{1}{2\sqrt{n\alpha}}; \theta^*\right) &\geq 1 - \alpha \end{aligned}$$

This means the random interval $(\hat{\theta}_n - \frac{1}{2\sqrt{n\alpha}}, \hat{\theta}_n + \frac{1}{2\sqrt{n\alpha}}; \theta^*)$ contains θ^* with probability $1 - \alpha$. We can note that the interval decreases as n increases, and increases as α decreases. *N.B.* $\hat{\theta}_n$ is a random variable, while θ^* is not.

Example 5.4 - Asymptotically Exact Confidence Interval when Flipping Coins

This is an improvement on the bound produced in Example 5.3.

Let $\mathbf{X} \stackrel{\text{iid}}{\sim} \text{Bernoulli}(\theta^*)$ for some $\theta^* \in [0, 1]$, $W \sim \text{Normal}(0, 1)$ and define $\hat{\theta}_n := \hat{\theta}_n(X_1, \dots, X_n)$. We shall show that

$$\frac{\sqrt{n}(\hat{\theta}_n - \theta^*)}{\sqrt{\hat{\theta}_n(1 - \hat{\theta}_n)}} \rightarrow_{\mathcal{D}} W$$

We know that $\text{Var}(X_1) = \theta^*(1 - \theta^*)$.

By the *Weak Law of Large Numbers* $\hat{\theta}_n \rightarrow_{\mathbb{P}} \theta^*$.

By the *Central Limit Theorem*

$$\frac{\sqrt{n}(\hat{\theta}_n - \theta^*)}{\sqrt{\hat{\theta}_n(1 - \hat{\theta}_n)}} \rightarrow_{\mathcal{D}} W$$

Define $Y_n = \frac{\sqrt{n}(\hat{\theta}_n - \theta^*)}{\sqrt{\theta^*(1 - \theta^*)}}$ and $Z_n = \frac{\sqrt{\theta^*(1 - \theta^*)}}{\sqrt{\hat{\theta}_n(1 - \hat{\theta}_n)}}$.

By the *Continuous Mapping Theorem* tells us that $Z_n \rightarrow_{\mathcal{D}} 1$ and $Z_n \rightarrow_{\mathbb{P}} 1$.

Hence, by *Slutsky's Theorem*

$$\frac{\sqrt{n}(\hat{\theta}_n - \theta^*)}{\sqrt{\hat{\theta}_n(1 - \hat{\theta}_n)}} = Y_n Z_n \rightarrow_{\mathcal{D}} W$$

This gives us random interval

$$\left(\hat{\theta}_n - z_{\alpha/2} \sqrt{\frac{\hat{\theta}_n(1 - \hat{\theta}_n)}{n}}, \hat{\theta}_n + z_{\alpha/2} \sqrt{\frac{\hat{\theta}_n(1 - \hat{\theta}_n)}{n}} \right)$$

This interval captures θ^* asymptotically (in n) with probability $1 - \alpha$.

N.B. $z_{\alpha} = \Phi^{-1}(1 - \alpha)$ where Φ is the cumulative density function of a $\text{Normal}(0, 1)$.

6 The Fisher Information

Remark 6.1 - Motivation

In the next part of the content we shall show that given $\mathbf{X}_n \stackrel{\text{iid}}{\sim} f(\cdot; \theta^*)$ then for sufficiently regular models

- i) There exists a lower bound on the achievable performance of any estimate of θ^* .
- ii) A scaled & centered sequence of maximum likelihood estimators $\{\hat{\theta}_n(\mathbf{X}_n)\}$ become asymptotically normal as $n \rightarrow \infty$.

Remark 6.2 - Measuring Performance of Estimator

We measure the performance of an estimator $\hat{\theta}$ in terms of variance, since its mean should be θ^* . Lower variance indicates better performance.

Definition 6.1 - The Score Function

Let $\ell(\theta; x) := \ln f(x; \theta)$.

The *Score Function* is a measure of the sensitivity of the likelihood function wrt θ

$$\ell'(\theta; x) := \frac{d}{d\theta} \ell(\theta; x) = \frac{\frac{d}{d\theta} f(x; \theta)}{f(x; \theta)} = \frac{L'(\theta; x)}{L(\theta; x)}$$

Remark 6.3 - θ^* is a turning point of $\ell(\theta; x)$

Note that under the *Fisher Information Regularity Conditions* we have that $\forall \theta \in \Theta$

$$\begin{aligned} \mathbb{E}(\ell'(\theta; X); \theta) &= \int_S \frac{\frac{d}{d\theta} f(x; \theta)}{f(x; \theta)} f(x; \theta) dx \\ &= \int_S \frac{d}{d\theta} f(x; \theta) dx \\ &= \frac{d}{d\theta} \int_S f(x; \theta) dx \\ &= \frac{d}{d\theta} (1) \\ &= 0 \end{aligned}$$

This shows that we expect the derivative to equal 0 at θ^* . Further, this means θ^* is a turning point of the log-likelihood function (hopefully a maximum).

Example 6.1 - Application of Remark 6.3

Let $X \sim \text{Poisson}(\theta)$. Then $f_X(x; \theta) = \frac{\theta^x}{x!} e^{-\theta} \mathbf{1}\{x \in \mathbb{N}\}$.

$$\begin{aligned} \implies \ell(\theta; x) &= -\theta + x \ln \theta - \ln x! \\ \implies \ell'(\theta; x) &= -1 + \frac{x}{\theta} \\ \implies \mathbb{E}(\ell'(\theta; X); \theta) &= -1 + \frac{\theta}{\theta} \\ &= 0 \end{aligned}$$

Definition 6.2 - Fisher Information Regularity Conditions

Let Θ be an open interval in \mathbb{R} and $f(x; \theta)$ be a pmf/pdf.

Below are conditions which a model is required to meet in order to be considered sufficiently regular such that *Fisher Information* can be drawn from it.

- i) Both $L'(\theta; x) = \frac{d}{d\theta} f(x; \theta)$ and $L''(\theta; x) = \frac{d^2}{d\theta^2} f(x; \theta)$ exist for any $x \in \mathcal{X}$.
- ii) $\forall \theta \in \Theta$ the set $S := \{x \in \mathcal{X} : f(x; \theta) > 0\}$ does not depend on $\theta \in \Theta$.
- iii) The identity below exists

$$\int_S \frac{d}{d\theta} f(x; \theta) dx = \frac{d}{d\theta} \int_S f(x; \theta) dx = 0$$

Definition 6.3 - Fisher Information

Fisher Information is a technique for measuring the amount of information that an observable random variable X carries about an unknown parameter θ upon which the probability of X depends.

Let $X \sim f(\cdots; \theta)$. Then the *Fisher Information* for any $\theta \in \Theta$ is

$$I(\theta) := \mathbb{E}(\ell'(\theta; X)^2; \theta) \geq 0$$

N.B. This is the *Expectation of the score, squared* \equiv *Second moment of the score*.

Remark 6.4 - Fisher Information

- i) *Fisher Information* is a function of the parameter, θ , not the data, X .
- ii) $I(\theta)$ can be thought of as being the average *information* brought by a single observation X about θ , assuming $X \sim f(\cdot; \theta)$.
- iii) Since $\forall \theta \in \Theta, \mathbb{E}(\ell'(\theta; X); \theta) = 0$ then

$$I(\theta) = \text{Var}(\ell'(\theta; X); \theta)$$

The variance of the score.

Example 6.2 - Fisher Information of Poisson

Let $X \sim \text{Poisson}(\theta)$.

From **Example 6.1** we know that $\ell'(\theta; x) = -1 + \frac{x}{\theta}$. Then

$$\begin{aligned} I(\theta) &= \text{Var}(\ell'(\theta; X); \theta) \\ &= \text{Var}\left(-1 + \frac{X}{\theta}; \theta\right) \\ &= \text{Var}\left(\frac{X}{\theta}; \theta\right) \\ &= \frac{1}{\theta^2} \text{Var}(X; \theta) \\ &= \frac{1}{\theta^2} \cdot \theta \text{ since } X \sim \text{Poisson}(\theta) \\ &= \frac{1}{\theta} \end{aligned}$$

Theorem 6.1 - Alternative Expression of Fisher Information

Let $f(x; \theta)$ be a pmf/pdf which satisfies the conditions of **Definition 6.2**. If

$$\forall \theta \in \Theta \quad \int_{\mathcal{X}} \frac{d^2}{d\theta^2} f(x; \theta) dx = \frac{d}{d\theta} \int_{\mathcal{X}} \frac{d}{d\theta} f(x; \theta) dx$$

Then

$$I(\theta) = -\mathbb{E}\left(\frac{d^2}{d\theta^2} \ell(\theta; X); \theta\right)$$

N.B. $\frac{d}{d\theta} \int_{\mathcal{X}} \frac{d}{d\theta} f(x; \theta) dx = 0$ by the regularity conditions.

Proof 6.1 - Theorem 6.1

By the *Quotient Rule*

$$\begin{aligned} \frac{d^2}{d\theta^2} \ell(\theta; x) &= \frac{d}{d\theta} \frac{\frac{d}{d\theta} f(x; \theta)}{f(x; \theta)} \\ &= \frac{\frac{d^2}{d\theta^2} f(x; \theta)}{f(x; \theta)} - \left(\frac{\frac{d}{d\theta} f(x; \theta)}{f(x; \theta)} \right)^2 \end{aligned}$$

Consequently

$$\begin{aligned} \mathbb{E}\left(\frac{d^2}{d\theta^2} \ell(\theta; X); \theta\right) &= \int_S \frac{\frac{d^2}{d\theta^2} f(x; \theta)}{f(x; \theta)} f(x; \theta) dx - \int_S \left(\frac{\frac{d}{d\theta} f(x; \theta)}{f(x; \theta)} \right)^2 f(x; \theta) dx \\ &= \int_S \frac{d^2}{d\theta^2} f(x; \theta) dx - \int_S \left(\frac{d}{d\theta} \ell'(\theta; x) \right)^2 f(x; \theta) dx \\ &= 0 - \mathbb{E}(\ell'(\theta; X)^2; \theta) \\ &= -I(\theta) \\ \Rightarrow I(\theta) &= -\mathbb{E}\left(\frac{d^2}{d\theta^2} \ell(\theta; X); \theta\right) \end{aligned}$$

7 Efficiency and The Cramer-Rao Bound

Definition 7.1 - IID Score Function

Let $\mathbf{X} \stackrel{\text{iid}}{\sim} f(\cdot; \theta)$ for some $\theta \in \Theta$. Then the *Score Function* is

$$\ell'_n(\theta; \mathbf{x}) := \frac{d}{d\theta} \ell_n(\theta; \mathbf{x}) \text{ where } \ell_n(\theta; \mathbf{x}) := \ln f_n(\mathbf{x}; \theta) = \sum_{i=1}^n \ell(\theta; x_i)$$

Definition 7.2 - IID Fisher Information

Let $\mathbf{X} \stackrel{\text{iid}}{\sim} f(\cdot; \theta)$ for some $\theta \in \Theta$. Then the *Fisher Information* is

$$I_n(\theta) := \mathbb{E}(\ell'_n(\theta; \mathbf{X})^2; \theta) = \text{Var}(\ell'_n(\theta; \mathbf{X}); \theta)$$

Theorem 7.1 - Relationship between IID Fisher Information & Fisher Information

Consider the situation where $\forall \theta \in \Theta, f_n(\mathbf{x}; \theta) = \prod_{i=1}^n f(x_i; \theta)$. Then

$$\forall \theta \in \Theta, I_n(\theta) = nI(\theta)$$

Proof 7.1 - Theorem 7.1

Let $\mathbf{X} \stackrel{\text{iid}}{\sim} f(\cdot; \theta)$. Then

$$\begin{aligned} I_n(\theta) &= \text{Var}(\ell'_n(\theta; \mathbf{X}); \theta) \\ &= \text{Var}\left(\sum_{i=1}^n \ell'(\theta; X_i); \theta\right) \\ &= n \text{Var}\left(\sum_{i=1}^n \ell'(\theta; X_1); \theta\right) \\ \implies I_n(\theta) &= nI(\theta) \end{aligned}$$

0 Appendix

Definition 0.1 - Gradient

$$\nabla f(\boldsymbol{\theta}; \mathbf{x}) := \left(\frac{\partial f(\boldsymbol{\theta}; \mathbf{x})}{\partial \theta_1}, \dots, \frac{\partial f(\boldsymbol{\theta}; \mathbf{x})}{\partial \theta_n} \right)$$

Definition 0.2 - Hessian

$$\nabla^2 f(\boldsymbol{\theta}; \mathbf{x}) := \begin{pmatrix} \frac{\partial^2 f(\boldsymbol{\theta}; \mathbf{x})}{\partial \theta_1^2} & \frac{\partial^2 f(\boldsymbol{\theta}; \mathbf{x})}{\partial \theta_1 \partial \theta_2} & \cdots & \frac{\partial^2 f(\boldsymbol{\theta}; \mathbf{x})}{\partial \theta_1 \partial \theta_n} \\ \frac{\partial^2 f(\boldsymbol{\theta}; \mathbf{x})}{\partial \theta_1 \partial \theta_2} & \frac{\partial^2 f(\boldsymbol{\theta}; \mathbf{x})}{\partial \theta_2^2} & \cdots & \frac{\partial^2 f(\boldsymbol{\theta}; \mathbf{x})}{\partial \theta_2 \partial \theta_n} \\ \vdots & \vdots & \ddots & \vdots \\ \frac{\partial^2 f(\boldsymbol{\theta}; \mathbf{x})}{\partial \theta_1 \partial \theta_n} & \frac{\partial^2 f(\boldsymbol{\theta}; \mathbf{x})}{\partial \theta_2 \partial \theta_n} & \cdots & \frac{\partial^2 f(\boldsymbol{\theta}; \mathbf{x})}{\partial \theta_n^2} \end{pmatrix}$$

0.1 Notation

Notation	Denotes
$Z_n \rightarrow_{\mathbb{P}} Z$	$\{Z_n\}_{n \in \mathbb{N}}$ converges in <i>Probability</i> to random variable Z .
$Z_n \rightarrow_{\mathcal{D}} Z$	$\{Z_n\}_{n \in \mathbb{N}}$ converges in <i>Distribution</i> to random variable Z .
$Z_n \rightarrow_{qm} Z$	$\{Z_n\}_{n \in \mathbb{N}}$ converges in <i>Quadratic Mean</i> to random variable Z .
$\theta \in \Theta \subseteq \mathbb{R}^{d_\theta}$	Scalar or vector parameter characterising a probability distribution
$\hat{\theta}$	Estimation for the value of the parameter θ
θ^*	True value of the parameter θ
\mathbb{P}	Probability measure $\mathbb{P} : \mathcal{F} \rightarrow [0, 1]$
Ω	Sample space
X	Scalar random variable
\mathcal{F}	Sigma field (Set of events)
χ	Support of rv X . A set χ is definitely in it <i>i.e.</i> $\mathbb{P}(X \in \chi; \theta) = 1$
\mathbf{X}	Vector consisting of scalar random variables

0.2 R

Command	Result
<code>hist(a)</code>	Plots a histogram of the values in array a
<code>mean(a)</code>	Returns the mean value of array a
<code>rbinom(s, n, p)</code>	Samples n of $Bi(n, p)$ random variables
<code>rep(v, n)</code>	Produces an array of size n where each entry has value v
<code>x <- v</code>	Maps value v to variable x

0.3 Probability Distributions

Definition 0.3 - Binomial Distribution

Let X be a discrete random variable modelled by a *Binomial Distribution* with n events and rate of success p .

$$p_X(k) = \binom{n}{k} p^k (1-p)^{n-k}$$

$$\mathbb{E}(X) = np \quad \& \quad \text{Var}(X) = np(1-p)$$

Definition 0.4 - Gamma Distribution

Let T be a continuous random variable modelled by a *Gamma Distribution* with shape parameter

α & scale parameter λ . Then

$$\begin{aligned} f_T(x) &= \frac{\lambda^\alpha x^{\alpha-1} e^{-\lambda x}}{\Gamma(\alpha)} \quad \text{for } x > 0 \\ \mathbb{E}(T) &= \frac{\alpha}{\lambda} \quad \& \quad \text{Var}(T) = \frac{\alpha}{\lambda^2} \end{aligned}$$

N.B. $\alpha, \lambda > 0$.

Definition 0.5 - Exponential Distribution

Let T be a continuous random variable modelled by a *Exponential Distribution* with parameter λ . Then

$$\begin{aligned} f_T(t) &= \mathbf{1}\{t \geq 0\} \cdot \lambda e^{-\lambda t} \\ F_T(t) &= \mathbf{1}\{t \geq 0\} \cdot (1 - e^{-\lambda t}) \\ \mathbb{E}(X) &= \frac{1}{\lambda} \quad \& \quad \text{Var}(X) = \frac{1}{\lambda^2} \end{aligned}$$

N.B. Exponential Distribution is used to model the wait time between decays of a radioactive source.

Definition 0.6 - Normal Distribution

Let X be a continuous random variable modelled by a *Normal Distribution* with mean μ & variance σ^2 .

Then

$$\begin{aligned} f_X(x) &= \frac{1}{\sqrt{2\pi\sigma^2}} e^{-\frac{(x-\mu)^2}{2\sigma^2}} \\ F_X(x) &= \frac{1}{\sqrt{2\pi\sigma^2}} \int_{-\infty}^x e^{-\frac{(y-\mu)^2}{2\sigma^2}} dy \\ M_X(\theta) &= e^{\mu\theta + \sigma^2\theta^2(1/2)} \\ \mathbb{E}(X) &= \mu \quad \& \quad \text{Var}(X) = \sigma^2 \end{aligned}$$

Definition 0.7 - Poisson Distribution

Let X be a discrete random variable modelled by a *Poisson Distribution* with parameter λ . Then

$$\begin{aligned} p_X(k) &= \frac{e^{-\lambda} \lambda^k}{k!} \quad \text{For } k \in \mathbb{N}_0 \\ \mathbb{E}(X) &= \lambda \quad \& \quad \text{Var}(X) = \lambda \end{aligned}$$

N.B. Poisson Distribution is used to model the number of radioactive decays in a time period.