

# Point estimation

Statistical Inference - PhD EcoStatData

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- This unit will cover the following **topics**:
  - Methods of finding estimators
  - Methods of evaluating estimators
  - Unbiasedness
  - Asymptotic evaluations
  - Robustness and model misspecification
- The rationale behind **point estimation** is quite simple:
- When sampling is from a **population** described by a pdf or a pmf  $f(\cdot; \theta)$ , knowledge of  $\theta$  yields knowledge of the entire population.
- Hence, it is natural to seek a method of finding a good **estimator** of the unknown point  $\theta$ .

# Methods of finding estimators

# Estimator

A **point estimator**  $\hat{\theta}$  is any function of the random sample  $Y_1, \dots, Y_n$ , namely

$$\hat{\theta}(\mathbf{Y}) = \hat{\theta}(Y_1, \dots, Y_n).$$

That is, any **statistic** is a point estimator.

- To streamline the presentation, we consider estimators that target the **unknown parameter**  $\theta$  rather than an arbitrary (non-one-to-one) **transformation**  $g(\theta)$ .
- Most theoretical results **extend naturally** to the general case  $g(\theta)$ .

- An **estimator**  $\hat{\theta}(Y_1, \dots, Y_n)$  is a function of the sample  $Y_1, \dots, Y_n$  and is a **random variable**.
- An **estimate**  $\hat{\theta}(y_1, \dots, y_n)$  is a function of the realized values  $y_1, \dots, y_n$  and is a **number**.
- We will write  $\hat{\theta}$  to denote both estimators and estimates whenever its meaning is clear from the context.

# Method of moments

- The **method of moments** is, perhaps, the oldest method of finding point estimators, dating back at least to Karl Pearson in the late 1800s.
- Let  $Y_1, \dots, Y_n$  be an iid sample from  $f(\cdot; \theta)$ ,  $\theta = (\theta_1, \dots, \theta_p)$ , and  $\Theta \subseteq \mathbb{R}^p$ . Moreover, define

$$m_r = \frac{1}{n} \sum_{i=1}^n Y_i^r, \quad \mu_r(\theta) = \mu_r(\theta_1, \dots, \theta_p) = \mathbb{E}_\theta(Y^r), \quad r = 1, \dots, p.$$

corresponding to the **population moment**  $\mu_r(\theta_1, \dots, \theta_p)$  and the **sample moment**  $m_r$ .

- The method of moments estimator  $\hat{\theta}$  is obtained by solving the following system of equations for  $(\theta_1, \dots, \theta_p)$ :

$$\mu_1(\theta_1, \dots, \theta_p) = m_1,$$

$$\mu_2(\theta_1, \dots, \theta_p) = m_2,$$

$$\vdots$$

$$\mu_p(\theta_1, \dots, \theta_p) = m_p.$$

- In general, it is **not guaranteed** that a **solution exists** nor its **uniqueness**.

# Asymptotic evaluation of the MM

- Moments estimators are not necessarily the best estimators, but **under reasonable conditions** they are **consistent**, they have converge rate  $\sqrt{n}$ , and they are **asymptotically normal**.
- Suppose  $(Y, Y^2, \dots, Y^p)$  has **covariance**  $\Sigma$ , then the multivariate **central limit theorem** implies that as  $n \rightarrow \infty$

$$\sqrt{n}\{m - \mu(\theta)\} \xrightarrow{d} Z, \quad Z \sim N_p(0, \Sigma),$$

where  $m = (m_1, \dots, m_p)$  and  $\mu(\theta) = (\mu_1(\theta), \dots, \mu_p(\theta))$ .

- Suppose also that  $\mu(\theta)$  is a **one-to-one** mapping and let  $g(\mu)$  be the inverse of  $\mu(\theta)$ , that is  $g = \mu^{-1}$ . We assume that  $g$  has **differentiable** components  $g_r(\cdot)$  for  $r = 1, \dots, p$ .
- The moments estimator can be written as  $\hat{\theta} = g(m)$  and  $\theta = g(\mu(\theta))$ . Then, as a consequence of the **delta method**, the following general result holds:

$$\sqrt{n}(\hat{\theta} - \theta) \xrightarrow{d} Z, \quad Z \sim N_p(0, D\Sigma D^T),$$

where  $D = [d_{rr'}]$  is a  $p \times p$  matrix whose entries are the derivatives  $d_{rr'} = \partial g_r(\mu)/\partial \mu_{r'}$ .

Refer to van der Vaart (1998), Theorem 4.1, pag. 35-36.

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## Example: beta distribution

- Let  $Y_1, \dots, Y_n$  be an iid random sample from a beta distribution of parameters  $\alpha, \beta > 0$  with density

$$f(y; \alpha, \beta) = \frac{\Gamma(\alpha + \beta)}{\Gamma(\alpha)\Gamma(\beta)} y^{\alpha-1} (1-y)^{\beta-1}, \quad 0 < y < 1.$$

- The **moment estimator** for  $(\alpha, \beta)$  is the (**explicitly available**) solution of the system of equations

$$m_1 = \frac{\alpha}{\alpha + \beta}, \quad m_2 = \frac{\alpha(\alpha + 1)}{(\alpha + \beta)(\alpha + \beta + 1)}.$$

- After some algebra we obtain the following relationship, which is a **smooth** and **regular** function of  $(m_1, m_2)$ :

$$\hat{\alpha} = m_1 \frac{m_1 - m_2}{m_2 - m_1^2}, \quad \hat{\beta} = (1 - m_1) \frac{m_1 - m_2}{m_2 - m_1^2}.$$

where  $\hat{\sigma}^2 = m_2 - m_1^2$  is the **sample variance**. **Remark**: is it possible that  $m_1 < m_2$ ?

## Example: beta distribution (food expenditure)

- We consider data on **proportion of income** spent on **food** for a random sample of 38 households in a large US city.

**Low** income level ( $n = 17$ )

```
[1] 0.07431 0.13548 0.08825 0.13728 0.09629 0.09160 0.13882 0.09670 0.10866
[10] 0.11629 0.18067 0.14539 0.15869 0.14910 0.09550 0.23066 0.14751
```

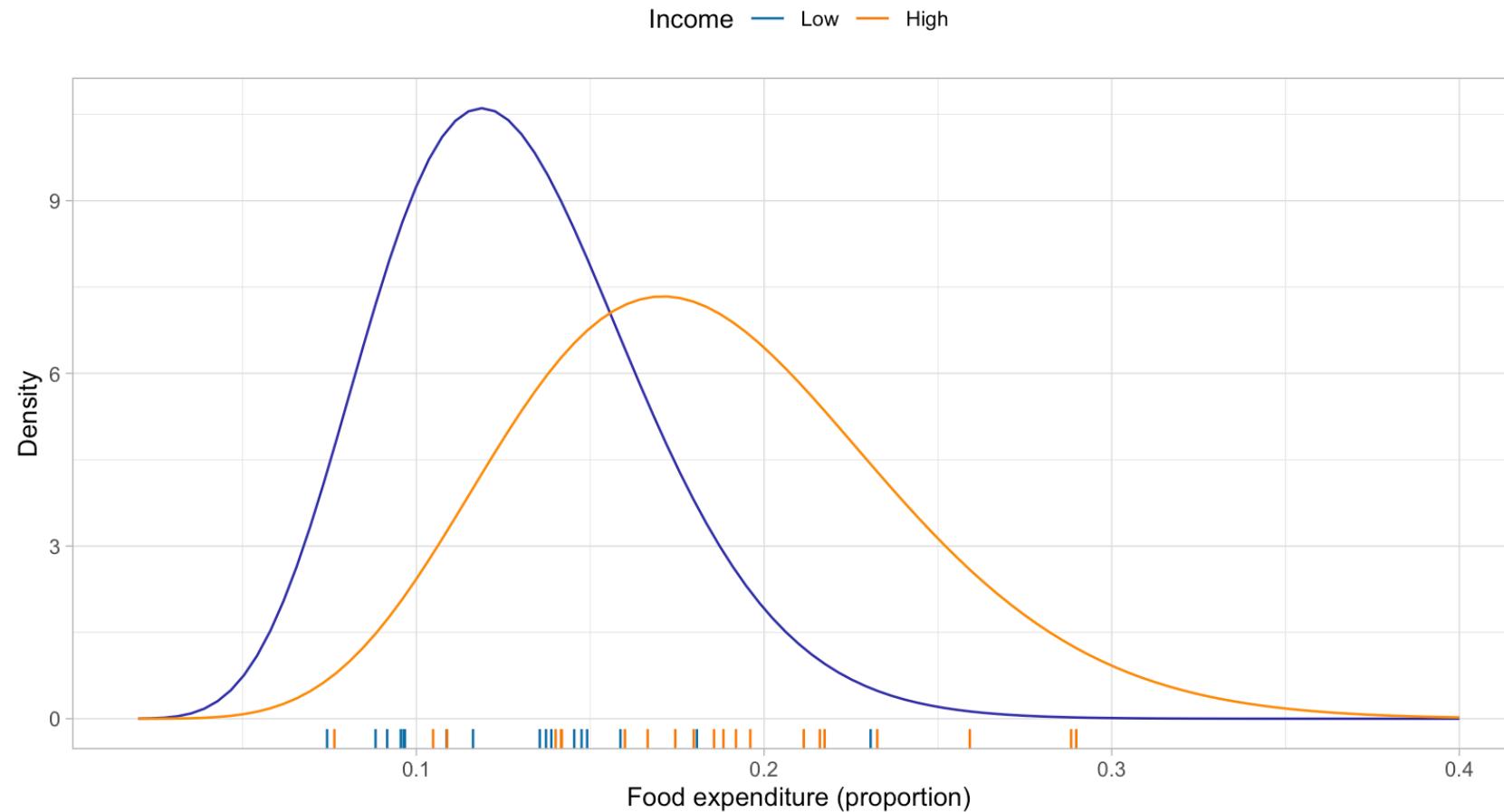
Here  $m_1 = 0.129$  and  $m_2 = 0.018$ , giving the **MM estimates**:  $\hat{\alpha} = 9.7$  and  $\hat{\beta} = 65.7$ .

**High** income level ( $n = 21$ )

```
[1] 0.15998 0.16652 0.21741 0.10481 0.23256 0.17976 0.14161 0.14184 0.19604
[10] 0.21141 0.17446 0.14005 0.18831 0.07641 0.21604 0.28980 0.10882 0.18561
[19] 0.19192 0.25918 0.28833
```

Here  $m_1 = 0.184$  and  $m_2 = 0.037$ , giving the **MM estimates**:  $\hat{\alpha} = 9$  and  $\hat{\beta} = 39.9$ .

## Example: beta distribution (food expenditure)



- **Estimated densities**  $f(y; \hat{\alpha}, \hat{\beta})$  for each income level, showing a reasonable fit in both cases.

## Example: binomial with unknown trials MM

- Let  $Y_1, \dots, Y_n$  be iid  $\text{Bin}(N, p)$  and we assume that **both**  $N$  and  $p$  are unknown.
- This is a somewhat unusual application of the binomial model, which can be used e.g. to estimate crime rates for crimes that are known to have many unreported occurrences.
- Equating the first two moments yields the system of equations

$$m_1 = Np, \quad m_2 = Np(1 - p) + N^2p^2.$$

- After some algebra we obtain the **moment estimator** for  $(N, p)$ , which is **smooth** and **regular** function of  $(m_1, m_2)$ :

$$\hat{p} = \frac{m_1}{\hat{N}}, \quad \hat{N} = \frac{m_1^2}{m_1 - \hat{\sigma}^2},$$

where  $\hat{\sigma}^2 = m_2 - m_1^2$  is the sample variance.

- Remark:** what if  $m_1 < \hat{\sigma}^2$ ?

This problem is described in Example 7.2.2 Casella and Berger (2002), pag. 313.

# Maximum likelihood estimator

- The method of **maximum likelihood** is, by far, the most popular technique for deriving estimators, developed by **Ronald A. Fisher** in Fisher (1922; 1925).
- Recall that  $L(\theta) = L(\theta; \mathbf{y})$  is the likelihood function and  $\ell(\theta) = \log L(\theta)$  is the log-likelihood.

Given a likelihood function  $L(\theta)$  of  $\theta \in \Theta$ , a **maximum likelihood estimate** of  $\theta$  is an element  $\hat{\theta} \in \Theta$  which attains the maximum value of  $L(\theta)$  in  $\Theta$ , i.e. such that  $L(\hat{\theta}) \geq L(\theta)$  or equivalently

$$L(\hat{\theta}) = \max_{\theta \in \Theta} L(\theta).$$

The **maximum likelihood estimator** (MLE) of the parameter  $\theta$  based on a sample  $\mathbf{Y}$  is  $\hat{\theta}(\mathbf{Y})$ .

- Intuitively, the MLE is a reasonable choice: it is the parameter point for which the observed sample is **most likely**.
- Clearly, the MLE is also the maximizer of the log-likelihood:  $\ell(\hat{\theta}) = \max_{\theta \in \Theta} \ell(\theta)$ .

# Properties and remarks about the MLE



- **Remark I:** the MLE may **not exists** and is **not** necessarily **unique**. On the other hand, if  $\Theta \subseteq \mathbb{R}^p$  and  $l(\theta)$  is differentiable, then it can be found as the solution of the **score equations**:

$$\ell^*(\theta) = \frac{\partial}{\partial \theta} l(\theta) = 0.$$

- **Remark II:** often  $\hat{\theta}$  cannot be written explicitly as a function of the sample values, i.e. in general the MLE has **no closed-form expression** but it must be found using **numerical procedures**.
- **Remark III:** the likelihood function has to be maximized in the set space  $\Theta$  specified by the statistical model, not over the set of the mathematically admissible values of  $\theta$ .

## Theorem (Invariance, Casella and Berger 2002, Theorem 7.2.10)

Let  $\psi(\cdot)$  be **one-to-one**<sup>1</sup>, that is, a **reparametrization**, from the set  $\Theta$  onto the set  $\Psi$ . Then the MLE of  $\psi = \psi(\theta)$  is  $\hat{\psi} = \psi(\hat{\theta})$  where  $\hat{\theta}$  denotes the MLE of  $\theta$ .

1. It generalizes to any  $g(\cdot)$ . If  $\hat{\theta}$  is the MLE, then  $g(\hat{\theta})$  is the “MLE” of an “induced likelihood”.

## Example: Poisson with unknown mean

- Let  $Y_1, \dots, Y_n$  be a iid random sample from a Poisson distribution of mean parameter  $\lambda > 0$ , with **likelihood function**

$$L(\lambda) = \prod_{i=1}^n \frac{e^{-\lambda} \lambda^{y_i}}{y_i!}.$$

- Therefore the **log-likelihood**, up to an additive constant  $c$  not depending on  $\lambda$ , is

$$\ell(\lambda) = \sum_{i=1}^n y_i \log \lambda - n\lambda + c.$$

- The **maximum likelihood estimator**  $\hat{\lambda}$  is found by maximizing  $\ell(\lambda)$ . In this regular problem, this can be done by studying the first derivative:

$$\ell^*(\lambda) = \frac{1}{\lambda} \sum_{i=1}^n y_i - n.$$

- We solve  $\ell^*(\lambda) = 0$ , obtaining  $\hat{\lambda} = \bar{y}$ . This is indeed a maximizer of  $\ell(\lambda)$  (**why?**).

## Example: binomial with unknown trials MLE

- Let  $Y_1, \dots, Y_n$  be iid  $\text{Bin}(N, p)$ , and suppose  $N$  is **unknown**, while  $p$  is considered **known**. This constitutes a non-regular problem because  $N$  is **discrete**.
- The likelihood function is

$$L(N) = \prod_{i=1}^n \binom{N}{y_i} p^{y_i} (1-p)^{N-y_i},$$

where the maximum **cannot** be obtained through **differentiation**, as  $N \in \mathbb{N}$ .

- Naturally, we require that  $\hat{N} \geq \max_i y_i$ , since  $L(N) = 0$  for any  $N < \max_i y_i$ . The ML is therefore an integer  $\hat{N} \geq \max_i y_i$  such that

$$L(\hat{N}) \geq L(\hat{N} - 1), \quad L(\hat{N} + 1) < L(\hat{N}).$$

- This value must be found **numerically**. However, it can be shown<sup>1</sup> that there exists exactly one such  $\hat{N}$ , meaning the MLE is **unique**.

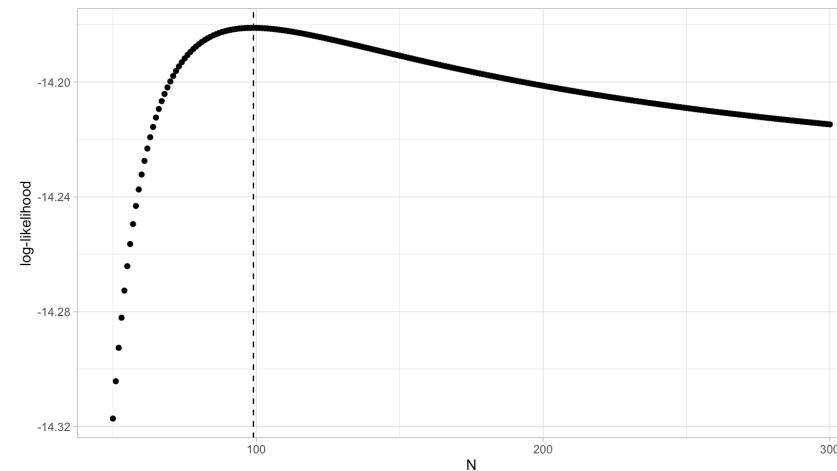
- This problem is described in Example 7.2.9 of Casella and Berger (2002), pag. 318. See also Example 7.2.13: such estimate has a large variance in practice.

## Example: binomial with unknown trials MLE

- Let us consider the following data, in which both  $N$  and  $p$  are unknown. These are **simulated data** and the **true values** were  $N = 75$  and  $p = 0.32$ .

```
[1] 16 18 22 25 27
```

- The **method of moments** estimator gives  $\hat{N}_{MM} = 102$  (rounded to the closest integer) and  $\hat{p}_{MM} = 0.21$ . The **maximum likelihood**, instead, gives  $\hat{N}_{ML} = 99$  and  $\hat{p}_{ML} = 0.22$ .



- If we replace 27 with a 28, we obtain drastically different estimates, namely  $\hat{N}_{MM} = 195$  and  $\hat{N}_{ML} = 191$ , demonstrating a **large** amount of **variability**.

## M-estimators

- M- and Z- estimators are broad class of estimators that encompass the maximum likelihood (iid observations) and other popular methods as special cases.<sup>1</sup>

An **M-estimator** is the **maximizer** over  $\Theta$  of a function  $M(\theta) : \Theta \rightarrow \mathbb{R}$  of the type:

$$M(\theta) = \sum_{i=1}^n m(\theta; Y_i),$$

where  $m(\theta; Y_i)$  are known real-valued functions.

- **Remark:** when  $m(\theta; y) = \log f(Y_i; \theta)$  this coincides with the MLE of a model with iid observations.

1. A detailed discussion is offered in van der Vaart (1998), Chap. 5.

# Z-estimators

A **Z-estimator** is the **solution** over  $\Theta$  of a system of equations function  $Q(\theta) = \mathbf{0}$  of the type:

$$Q(\theta) = Q(\theta; Y) = \sum_{i=1}^n q(\theta; Y_i) = \mathbf{0},$$

where  $q(\theta) = q(\theta; y)$  are known vector-valued maps. These are called **estimating equations**.

- When  $\theta = (\theta_1, \dots, \theta_p)$ ,  $Q$  and  $q$  typically have  $p$  coordinate functions, namely we consider:

$$Q_r(\theta) = \sum_{i=1}^n q_r(\theta; Y_i) = 0, \quad r = 1, \dots, p.$$

- In many examples  $q_r(y; \theta)$  are the partial derivatives of a function  $m(\theta; y)$ , that is

$$Q(\theta) = \frac{\partial}{\partial \theta} M(\theta).$$

An example is the **score function**  $\ell^*(\theta)$ . However, this is **not always the case**.

# Huber estimators I

- The **location** of a r.v.  $Y$  is a vague term that can be made precise by defining it as the expectation  $\mathbb{E}(Y)$ , a quantile, or the center of symmetry, as in the following example.
- Let  $Y_1, \dots, Y_n$  be a iid sample of real-valued random variables belonging to family of distributions  $\mathcal{F}$  defined as

$$\mathcal{F} = \{f(y - \theta) : \theta \in \Theta \subseteq \mathbb{R}\},$$

for some unknown density  $f(y)$  **symmetric** around 0. The parameter  $\theta$  is the **location**.

- Classical M- estimators for  $\theta$  are the **mean** and the **median**, maximizing:

$$-\sum_{i=1}^n (Y_i - \theta)^2, \quad (\text{Mean}) \qquad -\sum_{i=1}^n |Y_i - \theta|, \quad (\text{Median})$$

or alternatively (Z- estimator forms) solving the equations

$$\sum_{i=1}^n (Y_i - \theta) = 0, \quad (\text{Mean}) \qquad \sum_{i=1}^n \text{sign}(Y_i - \theta) = 0, \quad (\text{Median}).$$

## Huber estimators II

- **Huber estimators** can be regarded as a compromise between the mean and the median, maximizing the following function:

$$M(\theta) = - \sum_{i=1}^n m(Y_i - \theta), \quad m(y) = \begin{cases} \frac{1}{2}y^2 & \text{if } |y| \leq k \\ k|y| - \frac{1}{2}k^2 & \text{if } |y| \geq k \end{cases}$$

where  $k > 0$  is a **tuning** parameter. The function  $m(y)$  is continuous and differentiable<sup>1</sup>. The choice  $k \rightarrow 0$  leads to the median, whereas for  $k \rightarrow \infty$  we get the mean.

- Equivalently, we can consider the solution of the following estimating equation:

$$Q(\theta) = \sum_{i=1}^n q(Y_i - \theta) = 0, \quad q(y) = \begin{cases} -k & \text{if } y \leq -k \\ y & \text{if } |y| \leq k \\ k & \text{if } y \geq k \end{cases}$$

- Unfortunately, there is no closed-form expression and the equation must be solved numerically.

1. See Exercise 10.28 of Casella and Berger (2002).

## Example: Newcomb's speed of light

- Data  $y_1, \dots, y_{66}$  represent Simon Newcomb's measurements (1882) of the **speed of light**. The data are recorded as **deviations** from 24,800 nanoseconds.

```
[1] 28 26 33 24 34 -44 27 16 40 -2 29 22 24 21 25 30 23 29 31
[20] 19 24 20 36 32 36 28 25 21 28 29 37 25 28 26 30 32 36 26
[39] 30 22 36 23 27 27 28 27 31 27 26 33 26 32 32 24 39 28 24
[58] 25 32 25 29 27 28 29 16 23
```

- There are **two outliers** (-44 and -2) which could influence the analysis.
- We see that as  $k$  increases, the **Huber estimate** varies between the median (27) and the mean (26.21), so we interpret increasing  $k$  as decreasing **robustness** to outliers.
- The suggested **default** for  $k$  is roughly  $k \approx 4.5$ , that is  $k = 1.5 \times \text{MAD}$ .

$k$	0	5	10	20	30	40	50	60	70
Est.	27	27.37	27.417	27.125	26.831	26.677	26.523	26.369	26.215

- Based on recent measurements, the **true value** of the speed of light, expressed in this scale, is 33. **Huber estimate** for reasonable values of  $k$  is **closer** than both median and mean.

# Bayesian estimators

- Bayesian estimators are obtained following a **different inferential paradigm** than the one considered here, but they also exhibit **appealing frequentist properties**.
- Let  $L(\theta; \mathbf{y})$  denote the likelihood function, and let  $\pi(\theta)$  represent the **prior** distribution. Bayesian inference is based on the **posterior** distribution, defined as:

$$\pi(\theta | \mathbf{y}) = \frac{\pi(\theta)L(\theta; \mathbf{y})}{\int_{\Theta} \pi(\theta)L(\theta; \mathbf{y})d\theta}.$$

- Under certain hypotheses, which will be clarified later, the **posterior mean** serves as an **optimal Bayesian estimator**:

$$\hat{\theta}_{\text{Bayes}} = \mathbb{E}(\theta | \mathbf{Y}) = \frac{\int_{\Theta} \theta \pi(\theta)L(\theta; \mathbf{Y})d\theta}{\int_{\Theta} \pi(\theta)L(\theta; \mathbf{Y})d\theta}.$$

However, this estimator is not always available in closed form.

- Other “optimal” Bayesian estimators include, for instance, the posterior median.

## Example: binomial Bayes estimator

- Let  $Y_1, \dots, Y_n$  be iid Bernoulli random variables with probability  $p$ , and let the prior  $p \sim \text{Beta}(a, b)$ . Moreover, let  $n_1 = \sum_{i=1}^n y_i$  be the **number of successes** out of  $n$  trials.
- Standard calculations in Bayesian statistics yield the (**conjugate**) posterior  $(p \mid Y_1, \dots, Y_n) \sim \text{Beta}(a + n_1, b + n - n_1)$ . Hence, the **posterior mean** is

$$\hat{p} = \mathbb{E}(p \mid Y_1, \dots, Y_n) = \frac{a + n_1}{a + b + n}.$$

- Note that we can rewrite  $\hat{p}$  in the following way:

$$\hat{p} = \left( \frac{n}{a + b + n} \right) \bar{y} + \left( \frac{a + b}{a + b + n} \right) \left( \frac{a}{a + b} \right),$$

that is, as a **linear combination** of the prior mean and the sample mean, with weights determined by  $a, b$ , and  $n$ .<sup>1</sup>

- This is not a coincidence, and it essentially holds for general exponential families; see the elegant paper by Diaconis and Ylvisaker (1979).

# Methods of evaluating estimators

# Comparing estimators

- We study the performance of estimators, aiming to determine when an estimator  $\hat{\theta}$  can be considered **good** or **optimal**.
- This requires **criteria** for evaluation, often provided by **decision theory**, which relies on a **loss function**  $\mathcal{L}$ . The function  $\mathcal{L}(\theta, \hat{\theta})$  quantifies the loss incurred when estimating  $\theta$  by  $\hat{\theta}$ .
- Typically, we assume

$$\mathcal{L}(\theta, \theta) = 0 \text{ (no loss for the correct estimate)}, \quad \mathcal{L}(\theta, \hat{\theta}) \geq 0,$$

for all  $\theta$  and  $\hat{\theta}$ .

Since  $\hat{\theta}(\mathbf{Y})$  is random, we need a way to summarize the loss. A common criterion is the (frequentist) **risk function**, namely the **average loss**, defined as the expectation

$$R(\theta; \hat{\theta}) = \mathbb{E}_{\theta}\{\mathcal{L}(\theta, \hat{\theta}(\mathbf{Y}))\}.$$

# Optimal estimators

- An **oracle estimator**  $\hat{\theta}_{\text{oracle}}$ , which never makes errors, satisfies

$$R(\theta; \hat{\theta}_{\text{oracle}}) = 0,$$

for all  $\theta \in \Theta$ . Clearly, such an estimator **does not exist**.

- An **optimal estimator**  $\hat{\theta}_{\text{opt}}$  **uniformly** minimizes the risk, meaning

$$R(\theta; \hat{\theta}_{\text{opt}}) \leq R(\theta; \tilde{\theta}),$$

for all  $\theta \in \Theta$  and estimators  $\tilde{\theta}$ . Except for trivial cases, such an estimator **does not exist** unless one **restricts** the class of estimators considered <sup>1</sup>.

- In fact, the constant estimator  $\hat{\theta}_{\text{const}}(\mathbf{Y}) = \theta_1$  has zero risk when  $\theta = \theta_1$  but positive risk otherwise. Thus,  $\hat{\theta}_{\text{opt}}$  would need  $R(\theta; \hat{\theta}_{\text{opt}}) = 0$  for all  $\theta \in \Theta$ , making it identical to the oracle.

1. A common restriction is unbiasedness, namely the requirement  $\mathbb{E}_\theta(\hat{\theta}) = \theta$ . An alternative restriction is equivariance.

## Admissible estimators

An estimator  $\hat{\theta}$  is **admissible** if no other estimator  $\tilde{\theta}$  exists such that

$$R(\theta; \tilde{\theta}) \leq R(\theta; \hat{\theta}), \quad \text{for all } \theta \in \Theta,$$

with strict inequality for at least one value of  $\theta$ .

- Broadly speaking, an **admissible** estimator  $\hat{\theta}$  will perform better than an alternative  $\tilde{\theta}$  for some values of  $\theta$  and worse for others.
- In principle, disregarding any practical implications, an **inadmissible** estimator should not be used, as it is **dominated** by a better alternative.
- The admissibility criterion is interesting due to its **selective** nature, eliminating dominated estimators.
- However, **admissibility** alone is **not enough**: for instance, the constant estimator  $\hat{\theta}_{\text{const}} = \theta_1$  is admissible but clearly unsatisfactory.

# The choice of the loss function

- The choice of the loss function has its roots in **decision theory**. The loss function  $\mathcal{L}$  is a nonnegative function that generally increases as the distance between  $\hat{\theta}$  and  $\theta$  increases.
- Let  $\theta \in \mathbb{R}^p$ . Two **commonly used** loss functions are

$$\mathcal{L}(\theta, \hat{\theta}) = \|\hat{\theta} - \theta\|_1, \quad (\text{Absolute error loss}),$$

and

$$\mathcal{L}(\theta, \hat{\theta}) = \|\hat{\theta} - \theta\|_2^2, \quad (\text{Quadratic loss}).$$

- The quadratic loss is the de facto standard in many contexts, leading to the **mean squared error** and the well-known **bias-variance** decomposition. Let  $\Theta \subseteq \mathbb{R}$ , then

$$R(\theta; \hat{\theta}) = \mathbb{E}_\theta\{(\hat{\theta} - \theta)^2\} = \text{bias}_\theta(\hat{\theta})^2 + \text{var}_\theta(\hat{\theta}), \quad (\text{Mean squared error}).$$

- Both these losses are **convex**, and the quadratic loss is **strictly convex**. Convexity will be crucial in the subsequent theoretical developments.

# Other loss functions

- In principle, the choice of the loss should be based on its properties rather than mathematical convenience. Here we present some less common **examples**.
- **Weighted quadratic loss**, a variant of the quadratic loss, accounting for weights. Let  $\theta \in \mathbb{R}^p$

$$\mathcal{L}(\theta, \hat{\theta}) = (\theta - \hat{\theta})^T A (\theta - \hat{\theta}),$$

where  $A \in \mathbb{R}^{p \times p}$  is a positive definite matrix. This loss is **strictly convex**.

- **Stein Loss**<sup>1</sup>. Let  $\theta > 0$ , such as the **population variance** of a model. Stein loss is defined as:

$$\mathcal{L}(\theta, \hat{\theta}) = \frac{\hat{\theta}}{\theta} - 1 - \log \frac{\hat{\theta}}{\theta}.$$

A criticism of quadratic loss for variance estimation is that underestimation has finite penalty, while overestimation as infinite penalty. Instead, Stein loss  $\mathcal{L}(\theta, \hat{\theta}) \rightarrow \infty$  as  $\hat{\theta} \rightarrow 0$  and  $\hat{\theta} \rightarrow \infty$ .

- Other examples are **Huber losses** (see [Lehmann and Casella 1998, pp. 51–52](#)) and **intrinsic losses** (see [Robert 1994, p. 2.5.4](#)), such as the **entropy distance**.

1. See Examples 7.3.26 and 7.3.27 in Casella and Berger ([2002](#)) for a comparison in the estimation of the population variance  $\sigma^2$  under the quadratic loss and the Stein loss.

## Example: MSE of binomial estimators I

- Let  $Y_1, \dots, Y_n$  be iid Bernoulli random variables with probability  $p$ . Moreover, let  $n_1 = \sum_{i=1}^n y_i$  be the **number of successes** out of  $n$  trials.
- The **proportion**  $\hat{p} = \bar{y} = n_1/n$  is the **maximum likelihood** (and method of moments) estimator. Simple calculations yield

$$R(p; \hat{p}) = \mathbb{E}_p\{(\hat{p} - p)^2\} = \text{var}_p(\hat{p}) = \frac{p(1-p)}{n}.$$

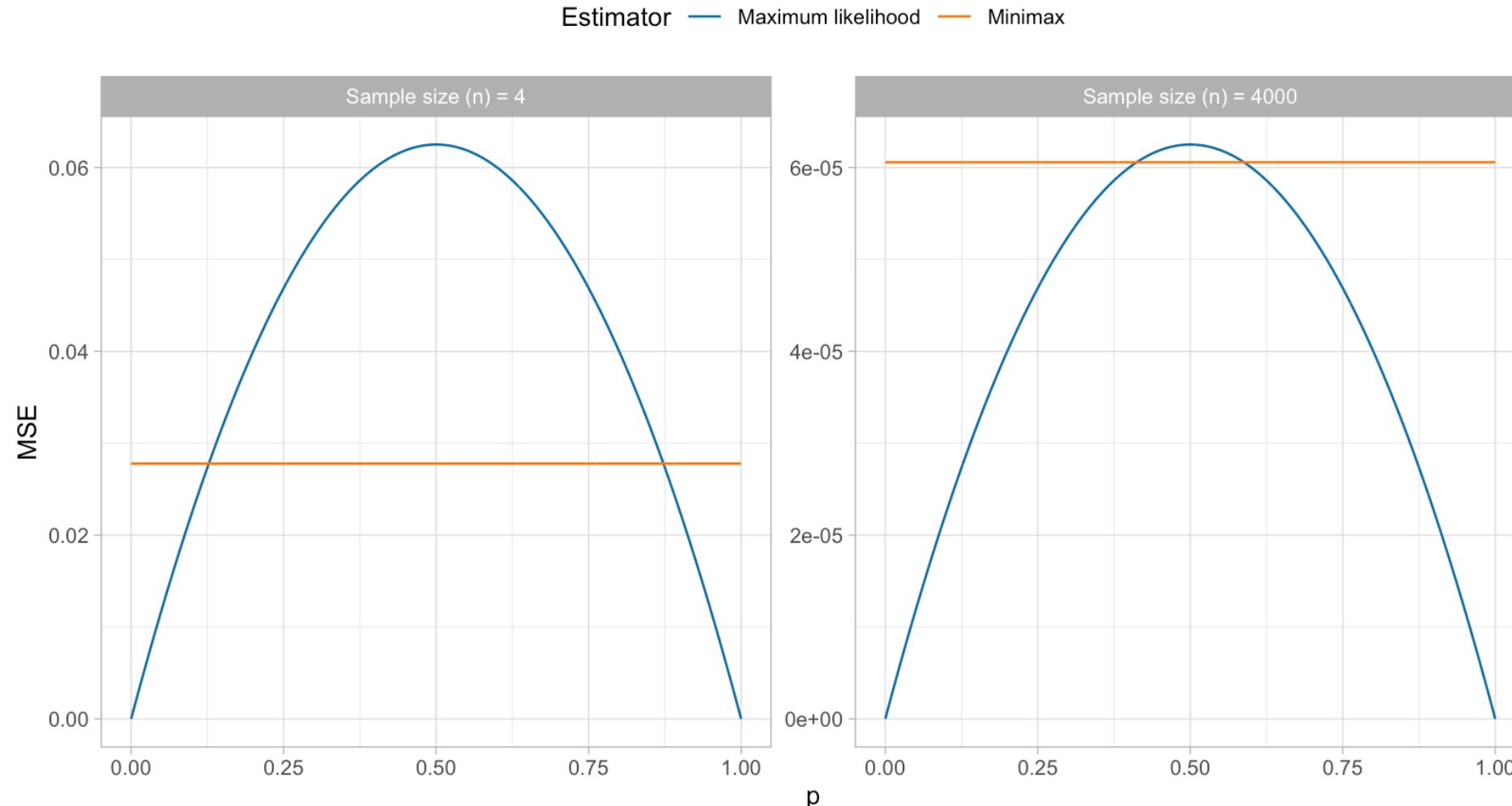
- Let us consider a **Bayesian estimator** for  $p$  under a beta prior with parameters  $a = b = 0.5\sqrt{n}$ , yielding

$$\hat{p}_{\text{minimax}} = \frac{n_1 + 0.5\sqrt{n}}{n + \sqrt{n}}, \quad R(p; \hat{p}_{\text{minimax}}) = \mathbb{E}_p\{(\hat{p}_{\text{minimax}} - p)^2\} = \frac{n}{4(n + \sqrt{n})^2}.$$

This Bayesian estimator has **constant risk**, that is,  $R(p; \hat{p}_{\text{minimax}})$  does not depend on  $p$ .<sup>1</sup>

- Some subjective Bayesians may criticize this estimator for not being “truly Bayesian” as the hyperparameters depend on the sample size  $n$ , making the prior data-dependent. However, here we are evaluating  $\hat{p}_{\text{Bayes}}$  from a frequentist perspective.

## Example: MSE of binomial estimators II



## Example: MSE of binomial estimators III

- Neither  $\hat{p}$  dominates  $\hat{p}_{\text{minimax}}$  nor vice versa. In fact, both estimators are **admissible**.
- For small values of  $n$ ,  $\hat{p}_{\text{minimax}}$  is the **better choice** unless one believes that  $p$  is very close to one of the extremes 0 or 1.
- Conversely, for large values of  $n$ , the maximum likelihood estimator  $\hat{p}$  is the better choice unless one strongly believes that  $p$  is very close to 0.5.
- This information, **combined** with the **knowledge of the problem** at hand, can lead to choosing the better estimator for the situation.

- We will get back to this example, as both  $\hat{p}$  and  $\hat{p}_{\text{minimax}}$  are **optimal** in some sense.
- In fact,  $\hat{p}$  is the **UMVU estimator** (uniform minimum variance unbiased estimator), and  $\hat{p}_{\text{minimax}}$  is the **minimax** estimator.

# James-Stein estimator I

- Let  $\mathbf{Y} = (Y_1, \dots, Y_p)$  be a **Gaussian random vector**,  $\mathbf{Y} \sim \mathcal{N}_p(\mu, I_p)$ , with unknown means  $\mu = (\mu_1, \dots, \mu_p)$  and fixed variance. That is,

$$Y_j \stackrel{\text{ind}}{\sim} \mathcal{N}(\mu_j, 1), \quad j = 1, \dots, p.$$

- We are interested in **estimating** the **means**  $\mu$ .

- Intuitively, the most “natural” approach is the maximum likelihood / method of moments estimator, which in this case with  $n = 1$  is simply

$$\hat{\mu}_j = Y_j, \quad j = 1, \dots, p.$$

- This estimator is “optimal” in the sense that it is the UMVUE, as we shall see in the next sections. Moreover, its **frequentist risk** under a squared loss is

$$R(\mu; \hat{\mu}) = \mathbb{E}_\mu(\|\hat{\mu} - \mu\|_2^2) = p,$$

because  $\|\hat{\mu} - \mu\|_2^2 = \|\mathbf{Y} - \mu\|_2^2 \sim \chi_p^2$ .

## James-Stein estimator II

- By the early 1950s, three proofs had emerged to show that  $\hat{\mu}$  is **admissible** for squared error loss **when**  $p = 1$ .
- Nevertheless, Stein (1956)<sup>1</sup> **stunned the statistical world** when he proved that although  $\hat{\mu}$  is admissible for squared error loss when  $p = 2$ , it is **inadmissible when**  $p \geq 3$ .
- In fact, James and Stein (1961)<sup>2</sup> showed that the estimator

$$\hat{\mu}_{JS} = \left(1 - \frac{p-2}{\|\mathbf{Y}\|_2^2}\right) \mathbf{Y}$$

strictly **dominates**  $\hat{\mu}$ .

1. Stein (1956), Inadmissibility of the usual estimator of the mean of a multivariate normal distribution, *Proc. Third Berkeley Symposium*, 1, 197–206, Univ. California Press.
2. James and Stein (1961), Estimation with quadratic loss, *Proc. Fourth Berkeley Symposium*, 1, 361–380, Univ. California Press.

## James-Stein estimator III

### Theorem (James-Stein, 1961)

Let  $\mathbf{Y} \sim \mathcal{N}_p(\mu, I_p)$ . The frequentist risk of the James-Stein estimator  $\hat{\mu}_{JS}$  under quadratic loss is

$$R(\mu; \hat{\mu}_{JS}) = \mathbb{E}_\mu(\|\hat{\mu}_{JS} - \mu\|_2^2) = p - (p-2)^2 \mathbb{E}_\mu \left( \frac{1}{\|\mathbf{Y}\|_2^2} \right).$$

Thus,  $\hat{\mu}_{JS}$  **strictly dominates**  $\hat{\mu}$ , meaning that  $R(\mu; \hat{\mu}_{JS}) < R(\mu; \hat{\mu}) = p$ . Moreover,

$$R(\mu; \hat{\mu}_{JS}) \leq p - \frac{(p-2)^2}{p-2 + \|\mu\|_2^2}.$$

- Geometrically, the James-Stein estimator **shrinks** each component of  $\mathbf{Y}$  towards the origin.
- The biggest **improvement** occurs when  $\mu$  is close to zero. For  $\mu = 0$ , we have  $R(0; \hat{\mu}_{JS}) = 2$  for all  $p \geq 2$ . As  $\|\mu\|_2^2 \rightarrow \infty$ , the risk approaches  $R(\mu; \hat{\mu}_{JS}) \rightarrow p$ .

## James-Stein estimator IV

- A useful **generalization** of the **James-Stein estimator** consists in **shrinking** the mean towards a **common value**  $m \in \mathbb{R}$  rather than 0, giving

$$\hat{\mu}_{\text{JS}}(m) = m + \left(1 - \frac{p-2}{\|\mathbf{Y} - m\|_2^2}\right) (\mathbf{Y} - m).$$

It holds that  $R(\mu; \hat{\mu}_{\text{JS}}(m)) \leq p - (p-2)^2/(p-2 + \|\mu - m\|_2^2)$ , therefore  $\hat{\mu}_{\text{JS}}(m)$  **dominates**  $\hat{\mu}$ .

- Even better, if we **estimate**  $m$  with the **arithmetic mean** of  $\mathbf{Y}$ , that is  $\bar{Y} = (1/p) \sum_{j=1}^p Y_j$ , we can obtain the following **shrinkage estimator**:

$$\hat{\mu}_{\text{shrink}} = \bar{Y} + \left(1 - \frac{p-3}{\|\mathbf{Y} - \bar{Y}\|_2^2}\right) (\mathbf{Y} - \bar{Y}).$$

Intuitively,  $p-3$  is the appropriate constant as an additional parameter is estimated. Moreover, in Efron and Morris (1975) it is proved that

$$R(\mu; \hat{\mu}_{\text{shrink}}) \leq p - \frac{(p-3)^2}{p-3 + \|\mu - \bar{\mu}\|_2^2},$$

with  $\bar{\mu} = (1/p) \sum_{j=1}^p \mu_j$ , again **dominating** the maximum likelihood  $\hat{\mu}$ .

## James-Stein estimator V

The James-Stein estimator is an empirical **Bayes estimator** in disguise. Let  $\mathbf{Y} \sim \mathbf{N}_p(\mu, I_p)$  and consider the **prior**  $\mu \sim N_p(m, \tau^2 I_p)$ . Then the **posterior mean** is

$$\hat{\mu}_{\text{Bayes}} = m + \left(1 - \frac{1}{1 + \tau^2}\right) (\mathbf{Y} - m).$$

James-Stein is an **empirical Bayes** approach: the quantity  $1/(1 + \tau^2)$  is **estimated**.

- In particular, under the Bayesian model

$$\|\mathbf{Y} - m\|_2^2 = \sum_{j=1}^p (Y_j - m)^2 \sim (1 + \tau^2) \chi_p^2$$

and therefore  $(p - 2)/(\|\mathbf{Y} - m\|_2^2)$  is an **unbiased estimate** for  $1/(1 + \tau^2)$ . In fact:

$$\mathbb{E} \left( \frac{p - 2}{\|\mathbf{Y} - m\|_2^2} \right) = \frac{1}{1 + \tau^2}.$$

- Alternative estimates for  $1/(1 + \tau^2)$  leads to **refined** James-Stein estimators.

# Efron and Morris (1975)

## Data Analysis Using Stein's Estimator and Its Generalizations

BRADLEY EFRON and CARL MORRIS\*

In 1961, James and Stein exhibited an estimator of the mean of a multivariate normal distribution having uniformly lower mean squared error than the maximum likelihood estimator. This result is now called the empirical Bayes context. Stein's rule and its generalizations are then applied to predict baseball averages, to estimate toxoplasmosis prevalence rates, and to estimate the exact size of Pearson's chi-square test with results from a computer simulation. In each of these examples, the mean square error of these rules is less than half that of the sample mean.

### 1. INTRODUCTION

Charles Stein [15] showed that it is possible to make a uniform improvement on the maximum likelihood estimator (MLE) in terms of total squared error risk when estimating several parameters from independent normal observations. Later James and Stein [13] presented a particularly simple estimator for which the improvement was quite substantial near the origin, if there are more than two parameters. This achievement leads immediately to a uniform, nontrivial improvement over the least squares (Gauss-Markov) estimators for the parameters in the usual formulation of the linear model. One might expect a rush of applications of this powerful new statistical weapon, but such has not been the case. Resistance has formed along several lines:

1. Misuse of the statistical interpretation of the mathematical formulation leading to Stein's result, in particular the sum of squared errors loss function;
2. Difficulties in adapting the James-Stein estimator to the situations that are inevitably arise in practice;
3. Long familiarity with the generally good performance of the MLE in applied problems;
4. A feeling that any gains possible from a "complicated" procedure like Stein's could not be worth the extra trouble.

(J. W. Tukey at the 1972 American Statistical Association meeting in New York City said that savings would not be more than ten percent in practical situations.)

We have written a series of articles [5, 6, 7, 8, 9, 10, 11] that cover Points 1 and 2. Our purpose here, and in a lengthier version of this report [12], is to illustrate the methods suggested in these articles on three applied problems and in that way deals with Points 3 and 4. Only one of the three problems, the toxoplasmosis data, is "real" in the sense of being generated outside the statistical world. The other two problems are contrived to illustrate in a realistic way the genuine difficulties and

rewards of procedures like Stein's. They have the added advantage of having the true parameter values available for comparison of methods. The examples chosen are the first and only ones considered for this report, and the favorable results typify our previous experience.

To review the James-Stein estimator in the simplest setting, suppose that for given  $\theta$ ,

$$X_i | \theta_i \stackrel{\text{ind}}{\sim} N(\theta_i, 1), \quad i = 1, \dots, k \geq 3, \quad (1.1)$$

meaning the  $|X_i|$  are independent and normally distributed with mean  $E_{\theta_i} X_i = \theta_i$  and variance  $\text{Var}_{\theta_i}(X_i) = 1$ . The example (1.1) typically occurs as a reduction to this canonical form from more complicated situations, as when  $X_i$  is a sample mean with known variance that is taken to be unity through an appropriate scale transformation. The unknown vector of means  $\theta = (\theta_1, \dots, \theta_k)$  is to be estimated with loss being the sum of squared component errors

$$L(\theta, \hat{\theta}) = \sum_{i=1}^k (\hat{\theta}_i - \theta_i)^2, \quad (1.2)$$

where  $\hat{\theta} = (\hat{\theta}_1, \dots, \hat{\theta}_k)$  is the estimate of  $\theta$ . The MLE, which is also the sample mean,  $\hat{\theta}^*(\mathbf{X}) = \mathbf{X} = (X_1, \dots, X_k)$  has constant risk  $k$ ,

$$R(\theta, \hat{\theta}^*) = E_{\theta} \sum_{i=1}^k (X_i - \theta_i)^2 = k, \quad (1.3)$$

$E_{\theta}$  indicating expectation over the distribution (1.1). James and Stein [13] introduced the estimator  $\hat{\theta}^*(\mathbf{X}) = (\delta_1^*(\mathbf{X}), \dots, \delta_k^*(\mathbf{X}))$  for  $k \geq 3$ ,

$$\delta_i^*(\mathbf{X}) = \mu_i + (1 - (k-2)/S)(X_i - \mu_i), \quad i = 1, \dots, k \quad (1.4)$$

with  $\mu = (\mu_1, \dots, \mu_k)$  any initial guess at  $\theta$  and  $S = \sum (X_i - \mu_i)^2$ . This estimator has risk

$$R(\theta, \hat{\theta}^*) = E_{\theta} \sum_{i=1}^k (\delta_i^*(\mathbf{X}) - \theta_i)^2 \quad (1.5)$$

$$\leq k - \frac{(k-2)^2}{k-2 + \sum (\theta_i - \mu_i)^2} < k, \quad (1.6)$$

being less than  $k$  for all  $\theta$ , and if  $\theta_i = \mu_i$  for all  $i$  the risk is two, comparing very favorably to  $k$  for the MLE.

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- Efron and Morris (1975, JASA) is a **classical paper** on the **practical relevance** of James-Stein's estimator.
- This approach was used in **sports analytics** to predict the batting averages of 18 major league players in 1970.
- As expected, **shrinkage estimators** significantly improve upon the maximum likelihood estimator.
- Stein's estimator was **3.50 times more efficient** than the MLE in this case.
- It also showcases a useful practical demonstration of **variance-stabilizing** transformations:
  - The original data are proportions, i.e. arguably not Gaussians.
  - After a suitable transformations they can be approximately regarded as normal.

# Efron and Morris (1975)

Name	At Bats $n_i$	Hits $x_i$	Mean $\hat{p}_i$	James-Stein	Rem. mean
Clemente	45	18	0.400	0.290	0.346
Robinson	45	17	0.378	0.286	0.298
Howard	45	16	0.356	0.282	0.276
Johnstone	45	15	0.333	0.277	0.222
...					
Williams	45	10	0.222	0.254	0.330
Campaneris	45	9	0.200	0.249	0.285
Munson	45	8	0.178	0.244	0.316
Alvis	45	7	0.156	0.239	0.200

- The batting averages  $\hat{p}_i$  out of  $n_i$  trials for  $n = 18$  players have been **transformed** via  $y_i = \sqrt{n_i} \arcsin(2p_i - 1)$ . The **James-Stein estimator** is applied to  $y_i$  and then transformed back.
- The **shrinkage effect** is evident and provides a better estimate of the **remaining batting average** for the season.

## James-Stein estimator VI

- James-Stein estimators were initially seen with suspicion:
  - How come that deliberately introducing **bias** improves the estimates?
  - How come that **learning from the experience of other** points can modify and even improve the individual estimates?
- These ideas are nowadays well established:
  - Modern **shrinkage estimators** such as **ridge regression** and **lasso** are widely used.
  - The notion of **borrowing of information** is at the heart of **random effects models**.
- The James-Stein theorem rigorously confirms the theoretical relevance of **indirect information**. Remarkably, this is based on **frequentist criteria** rather than Bayesian ones.
- A **simple proof** of the James-Stein theorem can be found in Efron (2010); see also Efron and Hastie (2016) for a modern and accessible perspective or **this divulgative article**.

## Criticism to the risk function approach

- Some authors, such as Robert (1994), have **criticized** the risk function approach for comparing estimators. The main arguments are the following:

- The frequentist paradigm evaluates estimators based on their **long-run performance**, without accounting for the given observations  $\mathbf{Y}$ . A client may wish for **optimal results** for the observed data, and **not someone else's data**.
- The risk function approach implicitly assumes the **repeatability** of the **experiment**, which has sparked controversy, especially among Bayesian statisticians.
- For instance, if new observations come to the statistician, she should make use of them, potentially modifying the way the experiment is conducted, as in medical trials.
- The risk function depends on  $\theta$ , preventing a **total ordering** of the estimators. Comparisons are difficult unless a **uniformly optimal** procedure exists, which is **rare**.

## Integrated risk

- Let us begin by finding a criterion that induces **total ordering** among estimators. A potential solution is taking the **average** risk function over values of  $\theta$ .

Let  $R(\theta, \hat{\theta}) = \mathbb{E}_\theta\{\mathcal{L}(\theta, \hat{\theta})\}$  be the **risk function**, and let  $\pi(d\theta)$  be a probability measure **weighting** the relevance of each  $\theta$ . Then

$$r(\pi, \hat{\theta}) = \int_{\Theta} R(\theta; \hat{\theta}) \pi(d\theta),$$

is called **integrated risk** or **Bayes risk**.

- Thus, an estimator  $\hat{\theta}$  is **preferable** over another  $\tilde{\theta}$  if  $r(\pi, \hat{\theta}) < r(\pi, \tilde{\theta})$ . Moreover, if  $\hat{\theta}$  **dominates**  $\tilde{\theta}$ , that is if  $\tilde{\theta}$  is **inadmissible**, then  $r(\pi, \hat{\theta}) < r(\pi, \tilde{\theta})$  for any choice of weights.

## Example: integrated risk of binomial estimators

- Let us consider the estimation of the **probability**  $p$  from a binomial experiment using the maximum likelihood  $\hat{p} = n_1/n$  and the Bayesian  $\hat{p}_{\text{minimax}} = (n_1 + 0.5\sqrt{n})/(n + \sqrt{n})$  estimators.
- We previously computed the **mean squared error** for both estimators. Let us now compute the integrated risk, assuming **uniform weights**  $\pi(dp) = dp$ , i.e., a **simple average** of the MSE.
- The **integrated risk** of the **maximum likelihood** estimator is

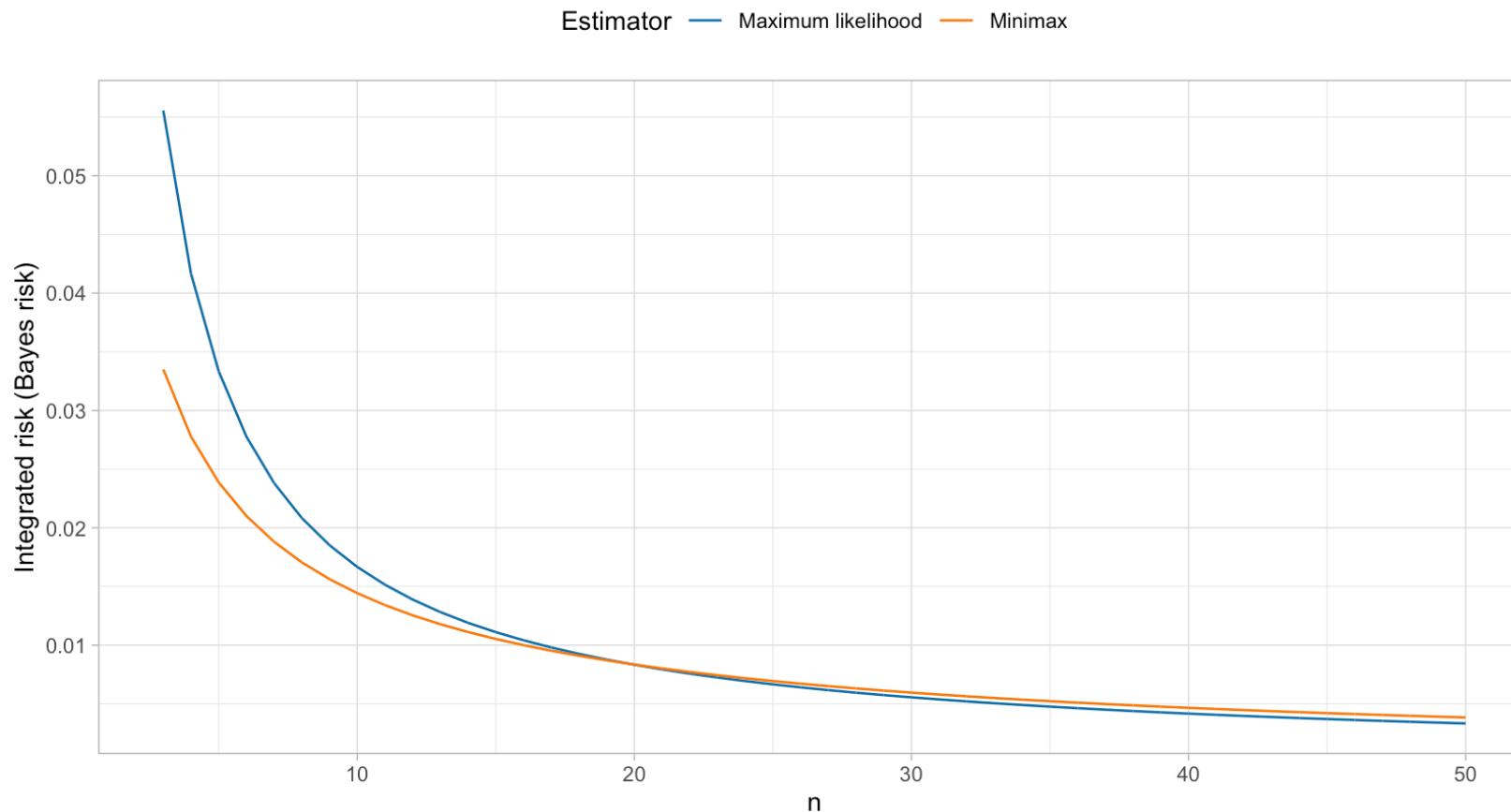
$$r(\pi, \hat{p}) = \int_0^1 \mathbb{E}_p\{(\hat{p} - p)^2\} dp = \frac{1}{n} \int_0^1 p(1-p) dp = \frac{1}{6n}.$$

- The **integrated risk** of the **Bayes estimator** coincides with the risk function, the latter being **constant** over  $p$ :

$$r(\pi, \hat{p}_{\text{minimax}}) = \int_0^1 \mathbb{E}_p\{(\hat{p}_{\text{minimax}} - p)^2\} dp = \int_0^1 \frac{n}{4(n + \sqrt{n})^2} dp = \frac{n}{4(n + \sqrt{n})^2}.$$

- Depending on the sample size  $n$ , one estimator may be preferable over the other.

# Example: integrated risk of binomial estimators



- According to the integrated risk and using constant weights  $\pi(dp) = dp$  one should **prefer**  $\hat{p}$  over  $\hat{p}_{\text{minimax}}$  when  $n \geq 20$  and viceversa.

# Bayesian estimators minimize the integrated risk



- The integrated risk is a sensible criterion for comparing estimators, provided a suitable set of weights  $\pi$  is selected. Hence, we may wish to find its **minimizer**.
- There is a surprisingly simple and elegant answer to this apparently difficult question.

**Theorem (Lehmann and Casella (1998), Theorem 1.1 in Chap. 4)**

Let  $\pi(d\theta)$  be the **prior distribution** for  $\theta$ . The **minimizer** of the **posterior expected loss**, if a unique solution exists, is called **Bayes estimator**. Moreover:

$$\hat{\theta}_{\text{Bayes}}(\mathbf{Y}) = \arg \min_{\tilde{\theta} \in \Theta} \int_{\Theta} \mathcal{L}(\theta, \tilde{\theta}(\mathbf{Y})) \pi(d\theta \mid \mathbf{Y}) = \arg \min_{\tilde{\theta} \in \Theta} r(\pi, \tilde{\theta}(\mathbf{Y})),$$

which means  $\hat{\theta}_{\text{Bayes}}$  coincides with minimizer of the **integrated risk**, provided  $r(\pi, \hat{\theta}_{\text{Bayes}}) < \infty$ .

- This fundamental theorem provides a **decision-theoretic justification** to Bayesian estimators as well as a practical recipe for finding them.

# Decision-theoretic justification of the posterior mean

**Corollary (Robert (1994), Proposition 2.5.1)**

Let  $\Theta \subseteq \mathbb{R}^p$ , then the **Bayes estimator** associated with prior distribution  $\pi$  and **quadratic loss**  $\mathcal{L}(\theta, \hat{\theta}) = \|\hat{\theta} - \theta\|_2^2$  is the **posterior mean**:

$$\hat{\theta}_{\text{Bayes}}(\mathbf{Y}) = \arg \min_{\tilde{\theta} \in \Theta} \int_{\Theta} \|\tilde{\theta}(\mathbf{Y}) - \theta\|_2^2 \pi(d\theta \mid \mathbf{Y}) = \mathbb{E}(\theta \mid \mathbf{Y}).$$

If the posterior mean exists, the Bayes estimator is **unique**.

- Hence, the **posterior mean** is **optimal** in the sense that minimizes the posterior expected loss and therefore also the integrated risk.
- This theorem extends to various loss functions. For instance if  $\mathcal{L}(\theta, \hat{\theta}) = \|\hat{\theta} - \theta\|_1$  then the **Bayes estimator** is the **posterior median**.

## Example: integrated risk of binomial estimators



- We previously considered the estimators  $\hat{p} = n_1/n$  and  $\hat{p}_{\text{minimax}} = (n_1 + 0.5\sqrt{n})/(n + \sqrt{n})$ . Assuming **uniform weights**, neither  $\hat{p}$  nor  $\hat{p}_{\text{minimax}}$  minimizes the integrated risk.
- In fact, theoretical results indicate that the **unique minimizer** is the **posterior mean** of  $p$  for a binomial model under a **uniform prior**  $\pi(dp) = dp$ . The optimal estimator is:

$$p \mid \mathbf{Y} \sim \text{Beta}(n_1 + 1, n - n_1 + 1) \implies \hat{p}_{\text{Bayes}} = \mathbb{E}(p \mid \mathbf{Y}) = \frac{n_1 + 1}{n + 2}.$$

- After some simple but tedious calculations, we obtain the associated **mean squared error**:

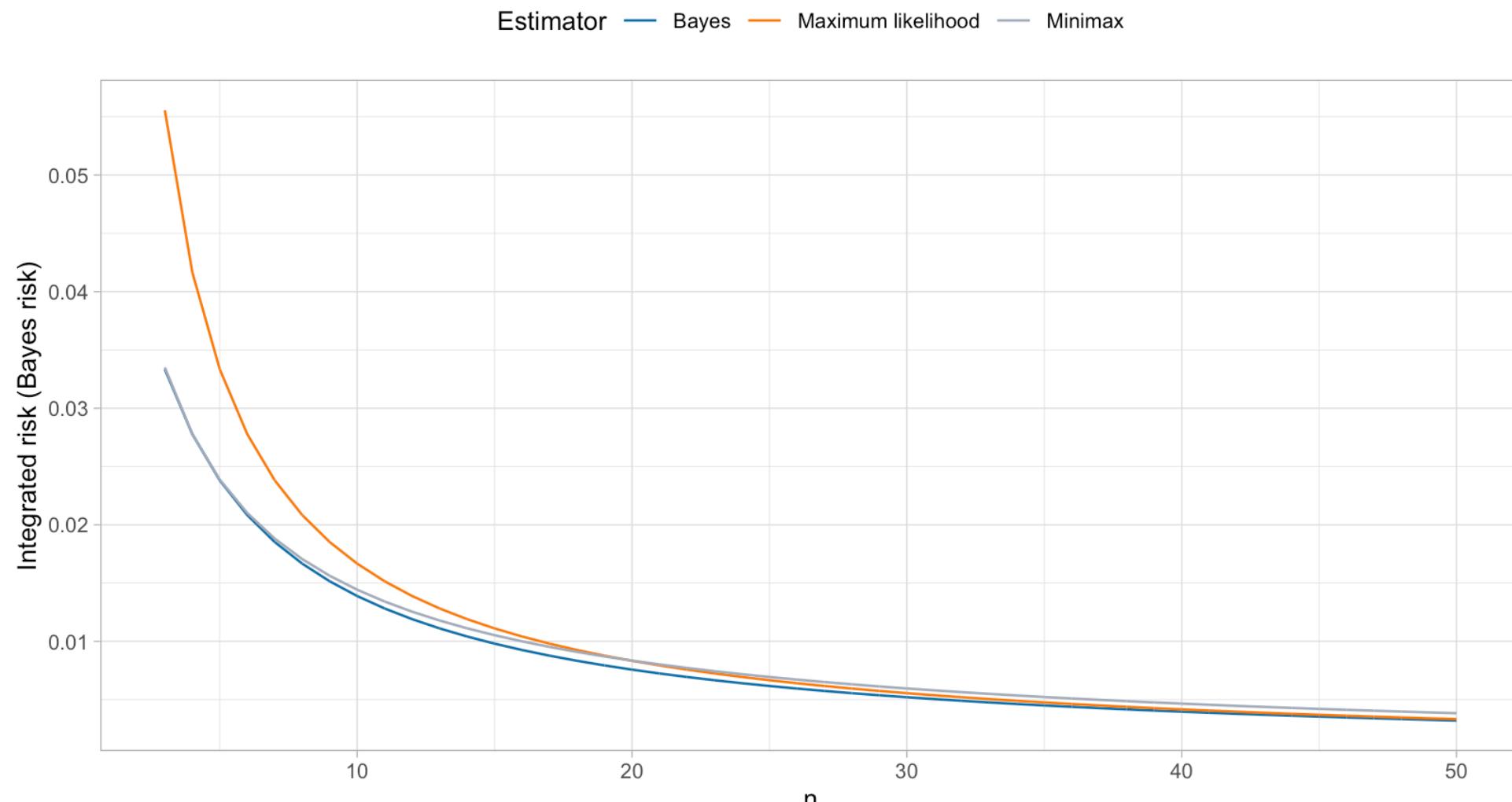
$$R(p; \hat{p}_{\text{Bayes}}) = \left( \frac{2}{n+2} \right)^2 (1/2 - p)^2 + \left( \frac{n}{n+2} \right)^2 \frac{p(1-p)}{n}.$$

- Integrating with respect to the **uniform prior** distribution, we obtain the **integrated risk**:

$$r(\pi, \hat{p}_{\text{Bayes}}) = \int_0^1 R(p; \hat{p}_{\text{Bayes}}) dp = \frac{1}{6(n+2)},$$

which is indeed **smaller** than  $r(\pi, \hat{p})$  and  $r(\pi, \hat{p}_{\text{minimax}})$ .

## Example: integrated risk of binomial estimators



# Admissibility of Bayesian estimators



- The next theorem provides a strong **frequentist justification** for **Bayesian estimators**.

**Theorem (Lehmann and Casella (1998), Theorem 2.4 in Chap. 5)**

Any unique Bayesian estimator is **admissible**.

- The uniqueness assumption is a technical condition often satisfied in practice. Under **squared loss**, or any other strictly convex loss, this holds automatically, provided the **estimator exists**.
- If the loss is strictly convex, such as the squared error loss, and the integrated risk is finite, this theorem remains valid even when using **improper priors** (Robert (1994), Proposition 2.4.25).
- **Admissibility** is a **minimum requirement**: even constant estimators are admissible.
- Nonetheless, the risk function approach dictates that inadmissible estimators should be discarded. This is non-trivial in practice, as evidenced by the James-Stein saga.

# Minimax estimators

- The **minimax** criterion is an alternative to the integrated risk for **discriminating** among (admissible) **estimators**. It comes from game theory, where two adversaries are competing.
- Instead of considering the “average” risk over  $\theta$  (integrated risk), the minimax criterion evaluates the **risk function** of estimators in the **worst-case scenario**.

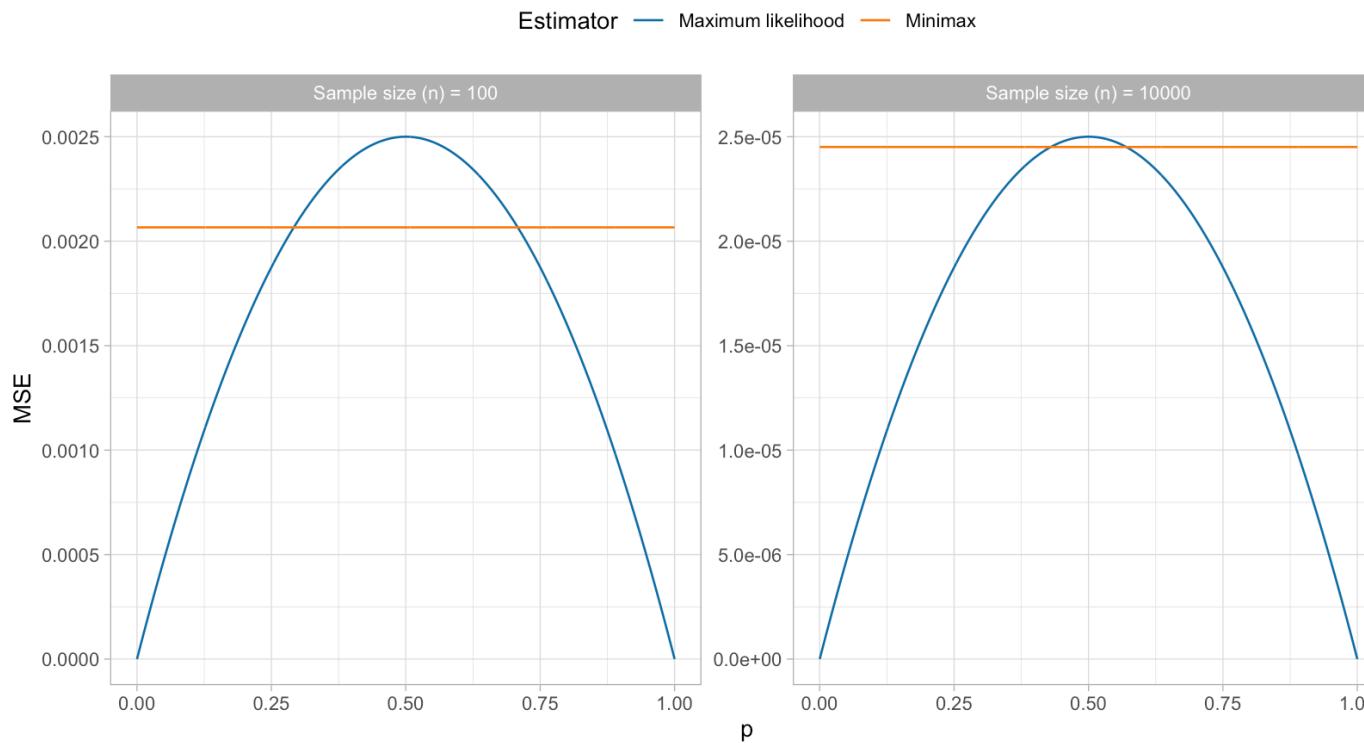
An estimator  $\hat{\theta}_{\text{minimax}}$  which **minimizes** the **maximum** risk, that is, which satisfies

$$\sup_{\theta \in \Theta} R(\theta; \hat{\theta}_{\text{minimax}}) = \inf_{\hat{\theta}} \sup_{\theta \in \Theta} R(\theta; \hat{\theta}),$$

is called **minimax estimator**. The quantity  $\sup_{\theta} R(\theta; \hat{\theta}_{\text{minimax}})$  is called **minimax risk**.

- In general, **finding minimax** estimators is **difficult**. Moreover, the resulting estimators are not necessarily very appealing in practice.

# Example: MSE of binomial estimators (minimax)



- We will show that  $\hat{p}_{\text{minimax}} = (n_1 + 0.5\sqrt{n})/(n + \sqrt{n})$  is indeed the **minimax estimator**. However, this is a very **conservative** choice.
- One could argue that the maximum likelihood  $\hat{p} = n_1/n$  is the preferred choice in practice, especially when  $n$  is large enough, even though it has slightly higher risk when  $p \approx 1/2$ .

# Minimax and Bayesian estimators

**Theorem (Lehmann and Casella (1998), Theorem 1.4, Chap. 5)**

Let  $\pi(d\theta)$  be a prior distribution and  $\hat{\theta}_{\text{Bayes}}$  be the **unique Bayes** estimator. If it holds

$$r(\pi, \hat{\theta}_{\text{Bayes}}) = \sup_{\theta} R(\theta; \hat{\theta}_{\text{Bayes}}),$$

then:

- the Bayes estimator  $\hat{\theta}_{\text{Bayes}}$  is also the **unique minimax** estimator.
- the prior  $\pi(d\theta)$  is **least favorable**, meaning that  $r(\pi, \hat{\theta}_{\text{Bayes}}) \geq r(\tilde{\pi}, \tilde{\theta}_{\text{Bayes}})$  for any other prior  $\tilde{\pi}(d\theta)$  and Bayes estimator  $\tilde{\theta}_{\text{Bayes}}$ .
- **Remark.** The above condition states the average of  $R(\theta; \hat{\theta}_{\text{Bayes}})$  is equal to its maximum. This will be the case when the risk function **constant** over  $\theta$ .
- More generally, if an **admissible estimator** has **constant risk**, is the unique **minimax** estimator (Robert (1994), Proposition 2.4.21).

# Unbiasedness

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## Unbiased estimators

An estimator  $\hat{\theta}$  is **unbiased** for  $\theta$  if  $\mathbb{E}_\theta(\hat{\theta}) = \theta$ , that is, if  $\text{bias}_\theta(\hat{\theta}) = \mathbb{E}_\theta(\hat{\theta} - \theta) = 0$  for all  $\theta \in \Theta$ .

- We often teach that **unbiasedness** is a natural and appealing property of an estimator. If  $\Theta \subseteq \mathbb{R}$  and under **squared error loss**, unbiasedness implies that

$$R(\theta; \hat{\theta}) = \mathbb{E}_\theta\{(\hat{\theta} - \theta)^2\} = \text{var}_\theta(\hat{\theta}).$$

- There are two main reasons for emphasizing unbiasedness:
  - It is often possible to find the uniformly “best” unbiased estimator, e.g., the one with the lowest variance (**UMVU estimator**).
  - For an estimator to be **consistent**, it must be at least **asymptotically unbiased**.

Unbiasedness is not a negative property per se. However, one may overlook better estimators by focusing too narrowly on this special class. Indeed, the **UMVUE** can even be **inadmissible**.

## Nonexistence of unbiased estimators

- Certain quantities **do not admit unbiased** estimators, even though they can be accurately estimated using **slightly biased** estimators.
- Let  $Y \sim \text{Bin}(n, p)$  and suppose we wish to find an estimator  $\hat{\psi}(Y)$  for the reparametrization  $\psi = 1/p$ . Then unbiasedness of an estimator  $\hat{\psi}(Y)$  would require

$$\sum_{k=0}^n \hat{\psi}(k) \binom{n}{k} p^k (1-p)^{n-k} = \frac{1}{p}, \quad p \in (0, 1).$$

Such an estimator **does not exist**!

- Indeed, the left hand side of this equation is a **polynomial**  $p$  with degree at most  $n$ . However,  $1/p$  cannot be written as a polynomial.
- Nonetheless, the **slightly biased** estimator  $\hat{\psi} = n/n_1$  will be close to  $1/p$  with high probability as  $n$  increases.

# Bayesian estimators and unbiasedness

**Theorem (Lehmann and Casella (1998), Theorem 2.3, Chap. 4)**

Let  $\Theta \subseteq \mathbb{R}^p$  and  $\hat{\theta}_{\text{Bayes}}(\mathbf{Y})$  be the unique Bayes estimator with prior  $\pi$  under a **squared error loss**. If  $\hat{\theta}_{\text{Bayes}}$  is **unbiased** then its integrated risk is

$$r(\pi, \hat{\theta}_{\text{Bayes}}) = 0.$$

- The above theorem is a formal way of saying that, apart from trivial cases, posterior means are **never unbiased** estimators.
- However, the **bias** comes with a **reduced variance**, therefore the trade-off could be favorable. This is guaranteed to occur because Bayesian estimators are admissible.
- Moreover, under mild regularity conditions, Bayesian estimators are **asymptotically unbiased**.

## Example: Poisson unbiased estimation

- Let  $Y_1, \dots, Y_n$  be i.i.d.  $\text{Poisson}(\lambda)$ , and let  $\bar{Y} = n^{-1} \sum_{i=1}^n Y_i$  and  $S^2 = (n-1)^{-1} \sum_{i=1}^n (Y_i - \bar{Y})^2$  be the **sample mean** and **sample variance**, respectively.
- It can be shown<sup>1</sup> that both estimators are **unbiased**, meaning

$$\mathbb{E}(\bar{Y}) = \mathbb{E}(S^2) = \lambda, \quad \text{for all } \lambda.$$

- To determine which estimator,  $\bar{Y}$  or  $S^2$ , is preferable, we should **compare their variances**. It is also well known that  $\text{var}_\lambda(\bar{Y}) = \lambda/n$ , whereas computing  $\text{var}_\lambda(S^2)$  can be **lengthy**.
- It holds that  $\text{var}_\lambda(\bar{Y}) < \text{var}_\lambda(S^2)$  for all  $\lambda$ . This implies that  $S^2$  is **inadmissible**.
- However, we can construct **infinitely** many **unbiased estimators** of  $\lambda$ :

$$\hat{\lambda}_a = a\bar{Y} + (1-a)S^2, \quad 0 < a < 1.$$

Is there a value of  $a$  such that  $\text{var}_\lambda(\hat{\lambda}_a) \leq \text{var}_\lambda(\bar{Y})$ ? What about other unbiased estimators?

- These are basic and well-known results: see Theorem 5.2.6 in Casella and Berger (2002).

## UMVU estimators

In this subsection on **unbiasedness**, we will often assume that  $\Theta \subseteq \mathbb{R}$ . All the results presented here extend to the **vector case**, though at the cost of heavier notation.

Let  $\Theta \subseteq \mathbb{R}$ . An estimator  $\hat{\theta}$  is a **best unbiased estimator** of  $\theta$  if it satisfies  $\mathbb{E}_\theta(\hat{\theta}) = \theta$  for all  $\theta$  (unbiasedness) and, for any **other unbiased** estimator  $\tilde{\theta}$ , we have

$$\text{var}_\theta(\hat{\theta}) \leq \text{var}_\theta(\tilde{\theta}), \quad \text{for all } \theta \in \Theta.$$

The estimator  $\hat{\theta}$  is also called **uniform minimum variance unbiased estimator** (UMVUE) of  $\theta$ .

- The UMVUE **does not necessarily exist**. If it does, **finding it** is not easy. And even if a unique UMVUE exists, it could still be **inadmissible**—recall the James-Stein saga.
- The success of UMVUE estimators is tied to two illuminating and elegant theorems: **Cramér-Rao** and **Rao-Blackwell**, which connect likelihood theory, sufficiency, and unbiasedness.

# Cramér-Rao inequality



- The Cramér-Rao theorem establishes a **lower bound** for the variance of an estimator. Thus, if the variance of an unbiased estimator  $\hat{\theta}$  attains the lower bound for all  $\theta$ , then  $\hat{\theta}$  is **UMVUE**.

## Theorem (Cramér-Rao, Theorem 7.3.9 in Casella and Berger (2002))

Let  $Y_1, \dots, Y_n$  be a sample from a joint probability measure  $f(\mathbf{y} \mid \theta)\nu(d\mathbf{y})$  and let  $\Theta \subseteq \mathbb{R}$ . Moreover, let  $\hat{\theta}(\mathbf{Y})$  be an estimator of  $\theta$  satisfying

$$1 + b^*(\theta) := \frac{\partial}{\partial \theta} \int \hat{\theta}(\mathbf{y}) f(\mathbf{y} \mid \theta) \nu(d\mathbf{y}) = \int \frac{\partial}{\partial \theta} \hat{\theta}(\mathbf{y}) f(\mathbf{y} \mid \theta) \nu(d\mathbf{y}),$$

and with finite variance  $\text{var}_\theta(\hat{\theta}(\mathbf{Y})) < \infty$ . Then

$$\text{var}_\theta(\hat{\theta}(\mathbf{Y})) \geq \frac{[1 + b^*(\theta)]^2}{\mathbb{E}_\theta(\ell^*(\theta)^2)}.$$

Moreover, if  $\hat{\theta}$  is an **unbiased** estimator for  $\theta$ , then  $b^*(\theta) = 0$  and  $\text{var}(\hat{\theta}(\mathbf{Y})) \geq 1/\mathbb{E}_\theta(\ell^*(\theta)^2)$ .

## Cramér-Rao: considerations

- The interchange of the derivative under the integral sign is an important **condition**, not merely a technical artifact of the proof.
- For example, if the **sample space** of i.i.d. random variables  $Y_i$  **depends on  $\theta$** , such condition is violated, and the Cramér-Rao lower bound may not hold.<sup>1</sup>
- The Cramér-Rao inequality is sometimes called **information inequality**. In fact  $I(\theta)$  defined as:

$$I(\theta) := \mathbb{E}_\theta(\ell^*(\theta)^2),$$

is called **Fisher information** or information number. This reflects the fact that as more information become available, the bound on the variance gets smaller.

If  $W(\mathbf{Y})$  is an unbiased estimator of a **transformation**  $g(\theta)$ , then the Cramér-Rao theorem holds as stated, but the term  $\frac{\partial}{\partial\theta}\mathbb{E}_\theta(W(\mathbf{Y})) = 1 + b^*(\theta)$  is not related to the “bias”.

1. See Example 7.3.13 in Casella and Berger (2002), for a simple and illuminating example.

# Bartlett identities I



## First Bartlett identity

Let  $Y_1, \dots, Y_n$  be a sample from a joint probability measure  $f(\mathbf{y} \mid \theta)\nu(d\mathbf{y})$  and let  $\Theta \subseteq \mathbb{R}$ . If we can interchange derivation and integration, namely

$$\frac{\partial}{\partial \theta} \int f(\mathbf{y} \mid \theta)\nu(d\mathbf{y}) = \int \frac{\partial}{\partial \theta} f(\mathbf{y} \mid \theta)\nu(d\mathbf{y}), \quad (\text{C.1})$$

then

$$\mathbb{E}_\theta(\ell^*(\theta)) = 0, \quad \text{implying} \quad I(\theta) = \mathbb{E}_\theta(\ell^*(\theta)^2) = \text{var}_\theta(\ell^*(\theta)).$$

- Thus, the regularity condition of Cramér-Rao implies that the score function  $\ell^*(\theta)$  is an **unbiased estimating equation**.

## Bartlett identities II

### Second Bartlett identity

Let  $Y_1, \dots, Y_n$  be a sample from a joint probability measure  $f(\mathbf{y} \mid \theta)\nu(d\mathbf{y})$  and let  $\Theta \subseteq \mathbb{R}$ . If we can interchange **twice** derivation and integration, namely

$$\frac{\partial}{\partial \theta} \int f(\mathbf{y} \mid \theta)\nu(d\mathbf{y}) = \int \frac{\partial}{\partial \theta} f(\mathbf{y} \mid \theta)\nu(d\mathbf{y}), \quad (\text{C.1})$$

$$\frac{\partial^2}{\partial^2 \theta} \int f(\mathbf{y} \mid \theta)\nu(d\mathbf{y}) = \int \frac{\partial^2}{\partial^2 \theta} f(\mathbf{y} \mid \theta)\nu(d\mathbf{y}), \quad (\text{C.2})$$

then

$$I(\theta) = \mathbb{E}_\theta(\ell^*(\theta)^2) = \text{var}_\theta(\ell^*(\theta)) = \mathbb{E}_\theta \left( -\frac{\partial^2}{\partial^2 \theta} \ell(\theta) \right).$$

- Both conditions are true in **regular exponential** families. This also clarifies that, in regular models, Fisher information relates to the **curvature** of the log-likelihood.

## Cramér-Rao: iid and regular case

### Theorem (Cramér-Rao, simplified)

Let  $Y_1, \dots, Y_n$  be an iid sample from  $f(y | \theta)dy$  satisfying conditions **C.1** and **C.2**, and let  $\Theta \subseteq \mathbb{R}$ . Moreover, let  $\hat{\theta}(\mathbf{Y})$  be an **unbiased** estimator of  $\theta$  with finite variance  $\text{var}(\hat{\theta}(\mathbf{Y})) < \infty$ . Then

$$\text{var}_\theta(\hat{\theta}(\mathbf{Y})) \geq \frac{1}{ni(\theta)}, \quad i(\theta) = - \int \left[ \frac{\partial^2}{\partial^2 \theta} \log f(y | \theta) \right] f(y | \theta) dy.$$

- The Fisher information is the **sum** of **individual contributions**  $I(\theta) = i(\theta) + \dots + i(\theta) = ni(\theta)$ .
- It can be shown<sup>1</sup> that **attainment**, that is the equality  $\text{var}(\hat{\theta}(\mathbf{Y})) = 1/ni(\theta)$ , occurs if and only if  $f(y | \theta)$  is the density of an **exponential family**.

1. See Theorem 5.12, Chap.2, Lehmann and Casella (1998).

## Example: Poisson unbiased estimation

- Let  $Y_1, \dots, Y_n$  be i.i.d.  $\text{Poisson}(\lambda)$ . The sample mean  $\bar{Y}$  is **unbiased** for  $\lambda$  and  $\text{var}_\lambda(\bar{Y}) = \lambda/n$ .
- We can use Cramér-Rao to show this estimator is a **UMVUE**. The regularity conditions are satisfied and therefore, after some calculus

$$i(\lambda) = -\mathbb{E}_\lambda \left( \frac{\partial}{\partial \lambda^2} \log f(y \mid \lambda) \right) = \mathbb{E}_\lambda \left( \frac{Y}{\lambda^2} \right) = \frac{1}{\lambda}.$$

- Hence, **Cramér-Rao** theorem states that for any unbiased estimator  $\hat{\lambda}(\mathbf{Y})$

$$\text{var}_\lambda(\hat{\lambda}) \geq \frac{\lambda}{n},$$

implying that  $\bar{Y}$  is a **UMVUE** because  $\text{var}_\lambda(\bar{Y}) = \lambda/n$ .

- Cramér-Rao theorem does not imply that  $\bar{Y}$  is the **unique** UMVUE.
- However, this is guaranteed by Theorem 7.3.19 in Casella and Berger (2002): if a **UMVUE** exists, it is **unique**.

## Cramér-Rao, multiparameter case

- The Cramér-Rao theorem naturally extends to the **vector case**, that is, when  $\Theta \subseteq \mathbb{R}^p$ .
- The **regularity conditions C.1** and **C.2** extend to the vector case, leading to the multiparameter **Bartlett identities**:

$$\mathbb{E}_\theta(\ell^*(\theta)) = \mathbf{0}, \quad I(\theta) = \mathbb{E}_\theta(\ell^*(\theta)\ell^*(\theta)^T) = \mathbb{E}_\theta\left(-\frac{\partial^2}{\partial\theta\partial\theta^T}\ell(\theta)\right),$$

where  $I(\theta)$  is called the **Fisher information matrix**, which is **positive definite**.

- Let  $Y_1, \dots, Y_n$  be an iid sample from  $f(y \mid \theta)dy$  satisfying the above regularity conditions. Moreover, let  $\hat{\theta}(\mathbf{Y})$  be an **unbiased** estimator of  $\theta$  with a finite covariance matrix. Then,

$$\text{var}_\theta(\hat{\theta}(\mathbf{Y})) \geq I(\theta)^{-1}, \quad I(\theta) = ni(\theta), \quad i(\theta) = -\int \frac{\partial^2}{\partial\theta\partial\theta^T} \log f(y \mid \theta)dy,$$

corresponding to the multiparameter Cramér-Rao theorem.

- Refer to Theorem 6.6, Chap. 2 in Lehmann and Casella (1998) for a proof.

# Rao-Blackwell

- The Rao-Blackwell theorem is **constructive strategy** for improving estimators that emphasizes the pivotal role of **sufficiency** in finding UMVU estimators.

**Theorem (Rao-Blackwell, Casella and Berger (2002), Theorem 7.3.17)**

Let  $\Theta \subseteq \mathbb{R}$  and  $\tilde{\theta}(\mathbf{Y})$  be an **unbiased** estimator of  $\theta$ . Moreover, let  $S = s(\mathbf{Y})$  be a **sufficient statistic** for  $\theta$  and  $\hat{\theta} = \mathbb{E}_{\theta}(\tilde{\theta}(\mathbf{Y}) \mid S)$ . Then  $\hat{\theta}$  is an estimator such that  $\mathbb{E}_{\theta}(\hat{\theta}) = \theta$  and

$$\text{var}_{\theta}(\hat{\theta}) \leq \text{var}_{\theta}(\tilde{\theta}).$$

That is,  $\hat{\theta}$  is a **uniformly better unbiased estimator** of  $\theta$ .

- As we shall see later, Rao-Blackwell holds for any **convex loss function**; see Lehmann and Casella (1998), Theorem 7.8, Chap. 1. Note that **unbiasedness** will not play any role.
- This means that conditioning of a sufficient statistic **reduces** the **mean squared error** of a biased estimator  $\tilde{\theta}$ , albeit the resulting  $\hat{\theta}$  would also be biased.

# En route to finding unique UMVUE I

- In looking for UMVUE we should only consider those based on a **sufficient statistic**  $S$ . However, if both  $\hat{\theta}$  and  $\tilde{\theta}$  are unbiased and based on  $S$ , how do we know if  $\hat{\theta}$  is best unbiased?
- The next theorem is a **partial answer**, which is useful if  $\hat{\theta}$  attains the **Cramér-Rao** lower bound.

**Theorem (Casella and Berger (2002), Theorem 7.3.19)**

Let  $\Theta \subseteq \mathbb{R}$ . If  $\hat{\theta}$  is a best unbiased estimator of  $\theta$ , then  $\hat{\theta}$  is **unique**.

**Example (Example 7.3.13 in Casella and Berger (2002))**

Let  $Y_1, \dots, Y_n$  be iid sample from a  $\text{Uniform}(0, \theta)$ . Then the estimator

$$\hat{\theta} = \frac{n+1}{n} \max\{Y_1, \dots, Y_n\}$$

is **unbiased** for  $\theta$  and is based on a **sufficient statistic**  $S = \max\{Y_1, \dots, Y_n\}$ . However, Cramér-Rao cannot be applied because the regularity conditions are not met. Is  $\hat{\theta}$  UMVUE?

## En route to finding unique UMVUE II

- Suppose we wish to improve on an unbiased estimator  $\hat{\theta}$ . Then, we could consider  $U = U(\mathbf{Y})$  such that  $\mathbb{E}_\theta(U) = 0$ , i.e.  $U$  is an **unbiased estimator or 0**, and let

$$\tilde{\theta} = \hat{\theta} + aU, \quad a \in \mathbb{R}.$$

Clearly,  $\tilde{\theta}$  is also **unbiased** and its **variance** is

$$\text{var}_\theta(\tilde{\theta}) = \text{var}_\theta(\hat{\theta}) + 2a\text{cov}_\theta(\hat{\theta}, U) + a^2\text{var}_\theta(U).$$

- If  $\text{cov}_\theta(\hat{\theta}, U) < 0$  for some  $\theta$ , then choosing  $a \in (0, -2\text{cov}_\theta(\hat{\theta}, U)/\text{var}_\theta(U))$  gives a **better estimator** for  $\theta$ , implying that  $\hat{\theta}$  is not UMVUE. This actually **characterizes** UMVU estimators.

**Theorem (Casella and Berger (2002), Theorem 7.3.20)**

Let  $\Theta \subseteq \mathbb{R}$  and  $\hat{\theta}$  be an unbiased estimator of  $\theta$ . Then  $\hat{\theta}$  is UMVUE **if and only if**  $\hat{\theta}$  is **uncorrelated** with all **unbiased estimators of 0**, that is

$$\text{cov}_\theta(\hat{\theta}, U) = 0, \quad \text{for all } U = U(\mathbf{Y}) \quad \text{such that } \mathbb{E}_\theta(U) = 0.$$

# Completeness

- Proving that an estimator  $\hat{\theta}$  is uncorrelated with all unbiased estimators of 0 is very hard, limiting the practical usefulness of the former theorem.
- However, if we assume  $S$  is **complete**, we can finally see the light at the end of the tunnel.

A sufficient statistic  $S = s(\mathbf{Y})$  for a statistical model  $\mathcal{F} = \{f(dy | \theta) : \theta \in \Theta\}$  is said to be **complete** if **no nonconstant** function  $g(\cdot)$  of  $S$  is **first-order ancillary**, that is,

$$\mathbb{E}_\theta(g(S)) = c \quad \text{for all } \theta \quad \text{implies} \quad g(S) = c \text{ a.s.}$$

In other words, any non-trivial transformation of  $S$  conveys information about  $\theta$  in expectation.

- Because of Rao-Blackwell, the UMVUE  $\hat{\theta} = \hat{\theta}(S)$  must be a function of a **sufficient** statistic  $S$ .
- If  $S$  is **complete**, then (using  $c = 0$ ), there exists no estimator  $U = U(S)$  such that  $\mathbb{E}_\theta(U(S)) = 0$ , with the only exception of  $U = 0$ , which is uncorrelated with  $\hat{\theta}(S)$ .

## Lehmann-Scheffé theorem

- We summarise these finding into a single statement, which is arguably the most relevant result of the Rao-Blackwell saga.

**Theorem (Lehmann-Scheffé-Rao-Blackwell, Casella and Berger (2002), Theorem 7.3.23)**

Let  $\Theta \subseteq \mathbb{R}$  and  $S$  be a **complete** sufficient statistic for a parameter  $\theta$ , and let  $\hat{\theta}$  be an unbiased estimator of  $\theta$  based only on  $S$ . Then,  $\hat{\theta}$  is **unique** best unbiased estimator (UMVUE) of  $\theta$ .

- The **Lehmann-Scheffé theorem** is implicitly contained in the previous results. Its **original formulation** is: “*unbiased estimators based on complete sufficient statistics are unique.*”
- The Lehmann-Scheffé theorem represents a **major achievement** in mathematical statistics, tying together **sufficiency**, **completeness** and **uniqueness**.

# Example: binomial best unbiased estimation

- Let  $Y_1, \dots, Y_n$  be iid  $\text{Bin}(N, p)$  and we assume that  $N$  is **known** and  $p$  is **unknown**. We are interested in estimating the **reparametrization**:

$$\psi = \mathbb{P}(Y_i = 1) = Np(1 - p)^{N-1}.$$

- We know that  $S = \sum_{i=1}^n Y_i \sim \text{Bin}(nN, p)$  is a **complete** and **sufficient** statistic<sup>1</sup>. However, no obvious unbiased estimator based on  $S$  is immediately available.
- Let us begin noting that  $\tilde{\psi} = \mathbb{I}(Y_1 = 1)$  is an **unbiased** estimator of  $\psi$ , where  $\mathbb{I}$  is the **indicator function**. In fact:  $\mathbb{E}(\tilde{\psi}) = \mathbb{E}(\mathbb{I}(Y_1 = 1)) = \mathbb{P}(Y_1 = 1) = \psi$ .
- The Lehmann-Scheffé-Rao-Blackwell theorem then implies that the unique and best unbiased estimator (UMVUE) is

$$\hat{\psi} = \mathbb{E}(\tilde{\psi} \mid S) = N \binom{Nn - N}{S - 1} / \binom{Nn}{S}.$$

which can be obtained after simple calculations.

- See Example 6.2.22 of Casella and Berger (2002). This also follows from general result of exponential families.

## Rao-Blackwell, multiparameter case

- Generalizations of Rao-Blakwell theory to the multiparameter case  $\Theta \subseteq \mathbb{R}^p$  are possible.
- If one is interested in estimating  $\psi = g(\theta)$  for some **function**  $g : \mathbb{R}^p \rightarrow \mathbb{R}$ , then the previously developed **theory applies almost directly**, with minor modifications to the statements.
- Note that a special case of the above is  $g(\theta) = \theta_j$ , meaning that the developed theory can be **separately** applied to **each coordinate** of  $\theta$ .

**Theorem (Rao-Blackwell, Lehmann and Casella (1998), Theorem 7.8, Chap. 1)**

Let  $\Theta \subseteq \mathbb{R}^p$ , and let  $\tilde{\theta}(\mathbf{Y})$  be an estimator of  $\theta$  with finite risk  $R(\theta; \tilde{\theta})$  under a **strictly convex loss** function. Moreover, let  $S = s(\mathbf{Y})$  be a **sufficient statistic** for  $\theta$ , and set  $\hat{\theta} = \mathbb{E}_\theta(\tilde{\theta}(\mathbf{Y}) \mid S)$ . Then,

$$R(\theta; \hat{\theta}) \leq R(\theta; \tilde{\theta}),$$

unless  $\tilde{\theta} = \hat{\theta}$  with probability 1.

# Alternative notions of optimality

# Unbiased estimating equations I

- A **Z-estimator** is the **solution** over  $\Theta$  of a system of equations function  $Q(\theta) = \mathbf{0}$  of the type:

$$Q(\theta) = \sum_{i=1}^n q(\theta; Y_i) = \mathbf{0},$$

where  $q(\theta) = q(\theta; y)$  are known vector-valued maps. These are called **estimating equations**.

The estimating equations  $Q(\theta) = \sum_{i=1}^n q(\theta; Y_i)$  are **unbiased** if they satisfy

$$\mathbb{E}_\theta(Q(\theta)) = \mathbf{0}, \quad \text{for all } \theta \in \Theta.$$

- Under iid sampling, the **score function** can be written as  $\ell^*(\theta) = \sum_{i=1}^n \frac{\partial}{\partial \theta} \log f(y_i; \theta)$  and therefore is a Z-estimator. Moreover, under regularity conditions, it is **unbiased**:  $\mathbb{E}_\theta(\ell^*(\theta)) = \mathbf{0}$ .

## Unbiased estimating equations II

- Remarkably, the unbiasedness of  $Q(\theta)$ , combined with a few regularity conditions, is often enough to prove **consistency** of the resulting estimator  $\hat{\theta}$  as  $n \rightarrow \infty$ .<sup>1</sup>
- Unbiasedness of  $Q(\theta)$  does not imply that the solution  $\hat{\theta}$  is an unbiased estimator of  $\theta$ , unless  $Q(\theta)$  is a **linear function**.
- The unbiasedness of  $Q(\theta)$  holds for any **reparametrization**  $\psi = \psi(\theta)$ . Moreover, Z-estimators, by construction, satisfy **equivariance**, meaning that  $\hat{\psi} = \psi(\hat{\theta})$ .

Having defined the class of unbiased estimating functions, the question naturally arises which of them we should use.

1. Refer to Davison (2003), Section 7.2, for an intuitive argument and van der Vaart (1998), Chap. 5 for a rigorous proof.

# Asymptotic behavior of unbiased estimating equations

- Guidance on the choice of  $Q(\theta)$  can be found by investigating its **asymptotic behavior**. We provide an **informal argument** for  $\theta \in \Theta \subseteq \mathbb{R}$ .
- Under **iid sampling** and further **regularity conditions**, implying that  $\mathbb{E}_\theta(q(\theta; Y_i)) = 0$ , a Taylor series expansion of  $Q(\theta)$  gives

$$0 = Q(\hat{\theta}) \approx \sum_{i=1}^n q(\theta; Y_i) + (\hat{\theta} - \theta) \sum_{i=1}^n \frac{\partial}{\partial \theta} q(\theta; Y_i).$$

- Thus, re-arranging, the following approximations hold **for large  $n$** :

$$\hat{\theta} - \theta \approx \frac{\sum_{i=1}^n q(\theta; Y_i)}{-\sum_{i=1}^n \frac{\partial}{\partial \theta} q(\theta; Y_i)} \approx \frac{Q(\theta)}{\mathbb{E}_\theta\left(-\frac{\partial}{\partial \theta} Q(\theta)\right)}.$$

- This suggests Z-estimators are **asymptotically unbiased** and the **asymptotic variance** of  $\hat{\theta}$  is

$$\text{var}(\hat{\theta}) \approx \frac{\text{var}(Q(\theta))}{\mathbb{E}_\theta\left(-\frac{\partial}{\partial \theta} Q(\theta)\right)^2} = n^{-1} \frac{\text{var}(q(\theta))}{\mathbb{E}_\theta\left(-\frac{\partial}{\partial \theta} q(\theta)\right)^2}.$$

## Godambe information

Let  $\mathbf{Y}$  be a sample from a statistical model with parameter  $\theta \in \Theta \subseteq \mathbb{R}$ . Let  $Q(\theta)$  be an **unbiased estimating equation**. We define the **sensitivity** as

$$H(\theta) := \mathbb{E}_\theta \left( -\frac{\partial}{\partial \theta} Q(\theta; \mathbf{Y}) \right),$$

and the **variability** as

$$J(\theta) := \text{var}_\theta(Q(\theta; \mathbf{Y})).$$

Then, the **Godambe information** is defined as

$$V(\theta) := \frac{H(\theta)^2}{J(\theta)}.$$

- An estimating equation  $Q(\theta)$  which has  $H(\theta) = J(\theta)$  is called **information unbiased**.

## Godambe efficiency

Let  $Q(\theta)$  and  $\tilde{Q}(\theta)$  be two **estimating equations** with Godambe information  $V(\theta)$  and  $\tilde{V}(\theta)$ . Then  $Q(\theta)$  is **uniformly more efficient** than  $\tilde{Q}(\theta)$  if

$$V(\theta) \geq \tilde{V}(\theta) \quad \text{for all } \theta \in \Theta \subseteq \mathbb{R}.$$

- This criterion is appropriate because the **inverse** of the **Godambe information**  $V(\theta)^{-1}$ , under regularity conditions, coincides with the **asymptotic variance** of  $\hat{\theta}$ .
- Moreover, although the **variance**  $J(\theta)$  is a natural basis for comparing estimating functions,  $\tilde{Q}(\theta) = aQ(\theta; \mathbf{Y})$  is also unbiased with variance  $\tilde{J}(\theta) = a^2 J(\theta)$ .
- Hence, a fair comparison is possible only after removing this **arbitrary scaling**. Indeed, Godambe information is **invariant** to scaling of  $Q(\theta)$ .

# Godambe information of the score function

- Let  $\mathbf{Y}$  be a sample from a statistical model with parameter  $\theta \in \Theta \subseteq \mathbb{R}$  and with **score function**  $\ell^*(\theta)$ .
- If the first Bartlett identity holds, the score function  $\ell^*(\theta)$  is an **unbiased estimating equation**.
- Moreover, if the second Bartlett identity holds,  $\ell^*(\theta)$  is **information unbiased**, that is

$$H(\theta) = \mathbb{E}_\theta \left( -\frac{\partial}{\partial \theta} \ell^*(\theta) \right) = \text{var}_\theta(\ell^*(\theta)) = J(\theta).$$

This implies that

$$V(\theta) = I(\theta),$$

that is, the **Godambe information** of  $\ell^*(\theta)$  coincides with the usual **Fisher information**.

- Hence, Godambe information is a generalization of Fisher information.

# Godambe efficiency

## Theorem (Godambe (1960))

Let  $Y_1, \dots, Y_n$  be a sample from a joint probability measure  $f(\mathbf{y} \mid \theta)\nu(d\mathbf{y})$  and let  $\Theta \subseteq \mathbb{R}$ .

Moreover, let  $Q(\theta)$  be an unbiased estimating equation with **Godambe information**  $V(\theta)$ . Then under **regularity conditions C.1** and **C.2**:

$$\frac{1}{V(\theta)} \geq \frac{1}{I(\theta)}.$$

- This result is the equivalent of the Cramér-Rao theorem for unbiased estimating equations.
- It implies that the **score functions** are **optimal** (Godambe efficient) among unbiased estimating equations. However, Z-estimators may have appealing **robustness** properties.

Godambe information generalizes to the vector case  $\theta \in \Theta \subseteq \mathbb{R}^p$  and is defined as  $V(\theta) = H(\theta)J(\theta)^{-1}H(\theta)$ , sometimes called **sandwich** matrix. Optimality results generalize as well.

# Linear estimating equations I

- Let  $Y_1, \dots, Y_n$  be **independent** random variables with **mean**  $\mathbb{E}(Y_i) = \mu_i(\theta)$  and **variance**  $\text{var}(Y_i) = V_i(\mu)$ , depending on a scalar parameter  $\theta \in \Theta \subseteq \mathbb{R}$ .
- Suppose the unbiased estimating equation  $Q(\theta)$  has a **linear form**:

$$Q(\theta) = \sum_{i=1}^n q(\theta; Y_i) = \sum_{i=1}^n w_i(\theta)(Y_i - \mu_i(\theta)),$$

for some set of positive **weights**  $w_i(\theta)$ . Can we find the **optimal** set of weights, according to the **Godambe information**?

- The **sensitivity**  $H(\theta)$  and the **variability**  $J(\theta)$  of  $Q(\theta)$  are readily available:

$$H(\theta) = \sum_{i=1}^n w_i(\theta)\mu_i^*(\theta), \quad J(\theta) = \sum_{i=1}^n w_i^2(\theta)V_i(\theta).$$

Therefore, the optimal weights are those that **maximize** the **Godambe information**:

$$V(\theta) = \frac{(\sum_{i=1}^n w_i(\theta)\mu_i(\theta))^2}{\sum_{i=1}^n w_i^2(\theta)V_i(\theta)}.$$

## Linear estimating equations II

- Using Lagrange multipliers, it can be shown<sup>1</sup> that the **optimal weights** are

$$w_i(\theta) \propto \frac{\mu_i^*(\theta)}{V_i(\theta)}, \quad \mu_i^*(\theta) = \frac{\partial}{\partial \theta} \mu(\theta), \quad i = 1, \dots, n,$$

which means the following unbiased **linear** estimating equation is **Godambe efficient**:

$$Q(\theta) = \sum_{i=1}^n \frac{\mu_i^*(\theta)}{V_i(\theta)} (Y_i - \mu_i(\theta)).$$

- This optimality property holds for a special class of **linear** unbiased estimating equations but does **not** make **assumptions** of the distribution of  $Y_i$  other than  $\mu_i(\theta)$  and  $V_i(\theta)$ .

Let  $Y_1, \dots, Y_n$  be independent random variables such that  $\mu_i(\theta) = \lambda e_i$  and  $V_i(\theta) = \lambda e_i$ , where  $e_i$  are **known constants** e.g. representing exposure. Then the optimal weights are  $w_i(\theta) = 1$  and  $\hat{\theta} = \bar{Y}/\bar{e}$ . Note the  $Y_i$  is **not** necessarily **Poisson**.

1. See Davison (2003), Section 7.2.2, for a proof.

# BLUE estimators I

**Theorem (Gauss-Markov, Agresti (2015), Section 2.7.1)**

Let  $\mathbf{Y} \in \mathbb{R}^n$  be a random vector satisfying  $\mathbb{E}(\mathbf{Y}) = \mathbf{X}\beta$  and  $\text{var}(\mathbf{Y}) = \sigma^2 I_p$ , where  $\mathbf{X}$  is an  $n \times p$  known matrix with **full rank**,  $\beta \in \mathbb{R}^p$  is an unknown vector, and  $\sigma^2 > 0$  is an unknown parameter.

Then the **best linear unbiased estimator** (BLUE) of  $\beta$  is

$$\hat{\beta} = (\mathbf{X}^T \mathbf{X})^{-1} \mathbf{X}^T \mathbf{Y}.$$

meaning that  $\hat{\beta}$  is **unbiased** and has the **minimum variance** among all unbiased linear estimators of  $\beta$ .

- The Gauss-Markov theorem does not make specific assumptions on the distribution of  $\mathbf{Y}$ , only on the **mean** and **variance**.
- If we strengthen the assumptions to  $\mathbf{Y} \sim N_p(\mathbf{X}\beta, \sigma^2 I_p)$ , then  $\hat{\beta}$  is the **UMVUE** of  $\beta$  among all estimators, including non-linear ones; see Lehmann and Casella (1998), Chap. 3, Sec. 4.
- As a special case of Gauss-Markov, if  $Y_i$  are iid with mean  $\mu$  and variance  $\sigma^2$ , then  $\bar{Y}$  is BLUE.

## BLUE estimators II

**Theorem (Aitken, Agresti (2015), Section 2.7.1)**

Let  $\mathbf{Y} \in \mathbb{R}^n$  be random vector satisfying  $\mathbb{E}(\mathbf{Y}) = \mathbf{X}\beta$  and  $\text{var}(\mathbf{Y}) = \Sigma$ , where  $\mathbf{X}$  is  $n \times p$  known matrix with **full rank**,  $\beta \in \mathbb{R}^p$  an unknown vector, and  $\Sigma$  an **known** covariance matrix. Then the **best linear unbiased estimator** (BLUE) of  $\beta$  is

$$\hat{\beta} = (\mathbf{X}^T \Sigma \mathbf{X})^{-1} \mathbf{X}^T \Sigma^{-1} \mathbf{Y}.$$

corresponding to the **generalized least squares** estimator.

- This interesting **generalization** of Gauss-Markov is often not applicable in practice, because the covariance matrix  $\Sigma$  is typically **unknown**.
- When  $\Sigma = \text{diag}(\sigma_1^2, \dots, \sigma_n^2)$  is **diagonal**, i.e. in presence of heteroschedasticity, the Aitken estimator reduces to the **weighted least squares** estimator.
- Moreover, if  $Y_i$  are independent random variables with mean  $\mu$  and variance  $\sigma_i^2$ , then the **weighted mean**  $\hat{\mu} = \sum_{i=1}^n w_i Y_i / \sum_{j=1}^n w_j$  is BLUE, where  $w_i = (1/\sigma_i^2)$ .

# BLUE estimators and unbiased estimating equations I

- We discuss here **connections** between the **BLUE estimators** and the **unbiased estimating equations**, aimed at providing a **unified view** of these concepts.
- Under the same assumptions of Gauss-Markov theorem, let us consider:

$$Q(\beta) = \mathbf{A}^T(\mathbf{Y} - \mathbf{X}\beta),$$

for some  $n \times p$  matrix  $\mathbf{A}$  having full rank. This is a **linear unbiased** estimating equation, and the **optimal** choice of  $\mathbf{A}$  maximizes the **Godambe information**.

- Solving this estimating equation we obtain to a linear unbiased estimator, which is

$$\hat{\beta} = (\mathbf{A}^T \mathbf{X})^{-1} \mathbf{A}^T \mathbf{Y}, \quad \mathbb{E}(\hat{\beta}) = (\mathbf{A}^T \mathbf{X})^{-1} \mathbf{A}^T \mathbf{X} \beta = \beta.$$

- The **sensitivity** and **variability** matrices of  $Q(\beta)$  are

$$J(\beta) = \text{var}(Q(\beta)) = \sigma^2 \mathbf{A}^T \mathbf{A}, \quad H(\beta) = \mathbb{E} \left( -\frac{\partial}{\partial \beta} Q(\beta) \right) = (\mathbf{A}^T \mathbf{X})^T.$$

Consequently, the **Godambe information** is  $V(\beta) = \sigma^{-2} (\mathbf{A}^T \mathbf{X})^T (\mathbf{A}^T \mathbf{A})^{-1} (\mathbf{A}^T \mathbf{X})^T$ .

## BLUE estimators and unbiased estimating equations II

- The **optimal** choice of  $\mathbf{A}$  equivalently minimizes the inverse of the **Godambe information**, which after a few algebraic manipulation is equal to

$$V(\beta)^{-1} = \sigma^2 \left[ (\mathbf{A}^T \mathbf{X})^{-1} \mathbf{A}^T \right] \left[ (\mathbf{A}^T \mathbf{X})^{-1} \mathbf{A}^T \right]^T.$$

- The key remark is the following: a direct calculation shows that the variance of  $\hat{\beta}$  is

$$\text{var}(\beta) = \sigma^2 \left[ (\mathbf{A}^T \mathbf{X})^{-1} \mathbf{A}^T \right] \left[ (\mathbf{A}^T \mathbf{X})^{-1} \mathbf{A}^T \right]^T = V(\beta)^{-1},$$

that is, the **variance** of  $\hat{\beta}$  coincides with the inverse of the **Godambe information**. This is a consequence of the linearity of  $Q(\beta)$ , otherwise the property holds only asymptotically.

- Thus, the same proof of Gauss-Markov theorem can be used to show that the **BLUE estimator** is **Godambe efficient** and the optimal matrix is  $\mathbf{A} = \mathbf{X}$ , giving

$$\hat{\beta} = (\mathbf{X}^T \mathbf{X})^{-1} \mathbf{X}^T \mathbf{Y}, \quad V(\beta) = \frac{1}{\sigma^2} \mathbf{X}^T \mathbf{X}.$$

- Moreover, if  $\mathbf{Y} \sim \mathcal{N}_p(\mathbf{X}\beta, \sigma^2 I_p)$ , then  $V(\beta)$  also coincides with the **Fisher information**  $I(\beta)$ .

# Asymptotic evaluations

# Asymptotic evaluations: preliminaries

- Asymptotic evaluations of estimators are taught in basic courses of inferential statistics. We focus on two main properties: **consistency** and **asymptotic normality**.
- An estimator  $\hat{\theta}_n$  is **consistent** if it converges **in probability** to the true value  $\theta_0$  as the sample size increases, i.e.,  $\hat{\theta}_n \xrightarrow{p} \theta_0$  as  $n \rightarrow \infty$ . Classical **sufficient conditions** are given below.

**Theorem (Casella and Berger (2002), Theorem 10.1.3)**

Let  $\Theta \subseteq \mathbb{R}$  and  $\hat{\theta}_n$  be a sequence of estimators such that the **asymptotic bias** and **variance** are zero, that is for every  $\theta \in \Theta$

$$\lim_{n \rightarrow \infty} \text{bias}_\theta(\hat{\theta}_n) = 0, \quad \lim_{n \rightarrow \infty} \text{var}_\theta(\hat{\theta}_n) = 0.$$

Then  $\hat{\theta}_n$  is a **consistent estimator** of  $\theta_0$ .

- Checking this condition case-by-case is difficult. Instead, we seek **general sufficient conditions** to establish the consistency of broad estimator classes, including **maximum likelihood**.

# Asymptotic evaluations: preliminaries

- The **asymptotic normality** of an estimator  $\hat{\theta}_n$  is a stronger property than consistency, implying that the estimator is **normally distributed** around the true value  $\theta_0$  for large  $n$ .
- Let  $\Theta \subseteq \mathbb{R}$  and let  $\hat{\theta}_n$  be a sequence of estimators. Under “regularity conditions”:

$$\sqrt{n}(\hat{\theta}_n - \theta_0) \xrightarrow{d} N(0, v(\theta_0)^{-1}), \quad n \rightarrow \infty,$$

where  $v(\theta_0)^{-1}$  is the so-called **asymptotic variance**.

- If  $\hat{\theta}_n$  is the **maximum likelihood**, then under regularity conditions  $nv(\theta)$  equals the **Fisher information**  $I(\theta) = ni(\theta)$  and the **asymptotic variance** is  $i(\theta_0)^{-1}$ .

In regular problems, the maximum likelihood estimator  $\hat{\theta}_n$  is **asymptotically efficient**, that is, for  $n$  large enough, its variance attains the Cramér-Rao lower bound. Informally,

$$\hat{\theta}_n \stackrel{\text{d}}{\sim} N(\theta_0, I(\theta_0)^{-1}).$$

- **Remark:** there exist infinitely many asymptotically efficient estimators.

## Example: Poisson with unknown mean

- Let  $Y_1, \dots, Y_n$  be a iid random sample from a **Poisson** distribution of mean parameter  $\lambda > 0$ . The maximum likelihood estimator  $\hat{\lambda}_n$  is the sample mean

$$\hat{\lambda}_n = \bar{Y}.$$

- One could invoke the strong **law of large numbers** to show that  $\hat{\lambda}_n \xrightarrow{\text{a.s.}} \lambda$ . Alternatively:

$$\mathbb{E}_\lambda(\hat{\lambda}) = \lambda, \quad \text{var}_\lambda(\hat{\lambda}_n) = \frac{\lambda}{n} = I(\lambda)^{-1},$$

which implies  $\text{bias}_\lambda(\hat{\lambda}_n) = 0$  and  $\lim_{n \rightarrow \infty} \text{var}_\lambda(\hat{\lambda}_n) = 0$ , from which **consistency** follows.

- Moreover, as a direct application of the **central limit theorem**, we also obtain that

$$\sqrt{n}(\hat{\lambda}_n - \lambda) \xrightarrow{\text{d}} \mathcal{N}(0, \lambda), \quad i(\lambda) = \frac{1}{\lambda}.$$

- In order to construct confidence intervals, one typically estimate the asymptotic variance  $i(\lambda)^{-1}$  with a consistent estimator  $i(\hat{\lambda}_n)^{-1} = \hat{\lambda}_n$ . Then **Slutsky theorem** ensures that

$$\sqrt{n} i(\hat{\lambda}_n)^{1/2}(\hat{\lambda}_n - \lambda) \xrightarrow{\text{d}} \mathcal{N}(0, 1).$$

# The classical “regularity conditions”

- (A1) We observe an **iid** sample  $Y_1, \dots, Y_n$  from a density  $f(y; \theta_0)$  with true value  $\theta_0 \in \Theta \subseteq \mathbb{R}^p$ .
- (A2) The model is **identifiable**, that is, the densities  $f(\cdot; \theta)$  and  $f(\cdot; \theta')$  are different for  $\theta \neq \theta'$ .
- (A3) The distributions  $f(y; \theta)$  have **common support**.
- (A4) The parameter space  $\Theta$  contains an **open set** of which  $\theta_0$  is an **interior point**.

We will sometimes need the additional conditions:

- (A5) There is a neighbourhood  $\mathcal{N}$  of the true value  $\theta_0$  within which the first **three derivatives** of  $\ell(\theta)$  exist a.s., and there exist functions  $m_{rst}(y)$  such that  $|\partial^3 f(y; \theta)/\partial\theta_r \partial\theta_s \partial\theta_t| \leq m_{rst}(y)$  and  $\mathbb{E}_\theta(M_{rst}(\mathbf{Y})) < \infty$  for  $r, s, t = 1, \dots, p$  and  $\theta \in \mathcal{N}$ .
- (A6) The **Fisher information matrix** is finite and **positive definite**, and

$$\mathbb{E}_\theta \left[ \frac{\partial}{\partial\theta_r} \ell(\theta) \right] = 0, \quad [I(\theta)]_{rs} = -\mathbb{E}_\theta \left[ \frac{\partial^2}{\partial\theta_r \partial\theta_s} \ell(\theta) \right] = \mathbb{E}_\theta \left[ \frac{\partial\ell(\theta)}{\partial\theta_r} \frac{\partial\ell(\theta)}{\partial\theta_s} \right],$$

for  $r, s = 1, \dots, p$ , that is, the **first** and **second Bartlett identities**.

# Wald inequality



**Theorem (Wald inequality, Lehmann and Casella (1998), Theorem 3.2, Chap. 6)**

Under assumptions **(A1)-(A3)**, by the strong law of large numbers:

$$\frac{1}{n} \ell(\theta; \mathbf{Y}) - \frac{1}{n} \ell(\theta_0; \mathbf{Y}) = \frac{1}{n} \sum_{i=1}^n \log \frac{f(Y_i; \theta)}{f(Y_i; \theta_0)} \xrightarrow{\text{a.s.}} -\text{KL}(f(\cdot; \theta_0) \mid f(\cdot; \theta)), \quad n \rightarrow \infty,$$

where **KL** is the **Kullback-Leibler divergence**, which is **strictly positive** for  $\theta \neq \theta_0$ . Thus:

$$\ell(\theta; \mathbf{Y}) - \ell(\theta_0; \mathbf{Y}) \xrightarrow{\text{a.s.}} -\infty.$$

Moreover, by the properties of the **KL**, or using Jensen's inequality, we deduce:

$$\mathbb{E}_{\theta_0} (\ell(\theta; \mathbf{Y})) < \mathbb{E}_{\theta_0} (\ell(\theta_0; \mathbf{Y})), \quad \theta \neq \theta_0,$$

which is commonly known as **Wald inequality**.

# Consistency for the MLE

- Wald inequality is the main workhorse for proving **consistency** of the **maximum likelihood**.
- Broadly speaking, **on average**, the likelihood is **maximized** at the true value  $\theta_0$ , and this holds almost surely for large  $n$ , suggesting that  $\hat{\theta}_n \rightarrow \theta_0$  almost surely as  $n \rightarrow \infty$ .
- We may be tempted to conclude that conditions **(A1)-(A3)** are enough to prove consistency. Unfortunately, there are **complications** on general spaces  $\Theta$ . An exception is given below:

**Theorem (Lehmann and Casella (1998), Cor. 3.5, Chap. 6)**

Under assumptions **(A1)-(A3)** and if  $\Theta = (\theta_0, \theta_1, \dots, \theta_k)$  is **finite**, then  $\hat{\theta}_n$  exists, is unique and is a **consistent estimator** of  $\theta_0$ .

**Proof.** Let  $A_{j,n} = \{\ell(\theta_0) > \ell(\theta_j)\}$  for  $j = 1, \dots, k$ . It holds  $\mathbb{P}(A_{j,n}) \rightarrow 1$  as  $n \rightarrow \infty$ . Hence:

$$\mathbb{P}(\ell(\theta_0) > \ell(\theta_j) \text{ for all } j = 1, \dots, k) = \mathbb{P}(\bigcap_{j=1}^k A_{j,n}) \geq 1 - \sum_{j=1}^k \mathbb{P}(A_{j,n}^C) \rightarrow 1, \quad n \rightarrow \infty.$$

# What could go wrong?

- Let  $Y_1, \dots, Y_n$  be an iid sample from the following univariate density<sup>1</sup>

$$f(y; \theta) = \frac{1}{2}\phi(y; 0, 1) + \frac{1}{2}\phi(y; \theta, e^{-2/\theta^2}), \quad \theta \in \mathbb{R},$$

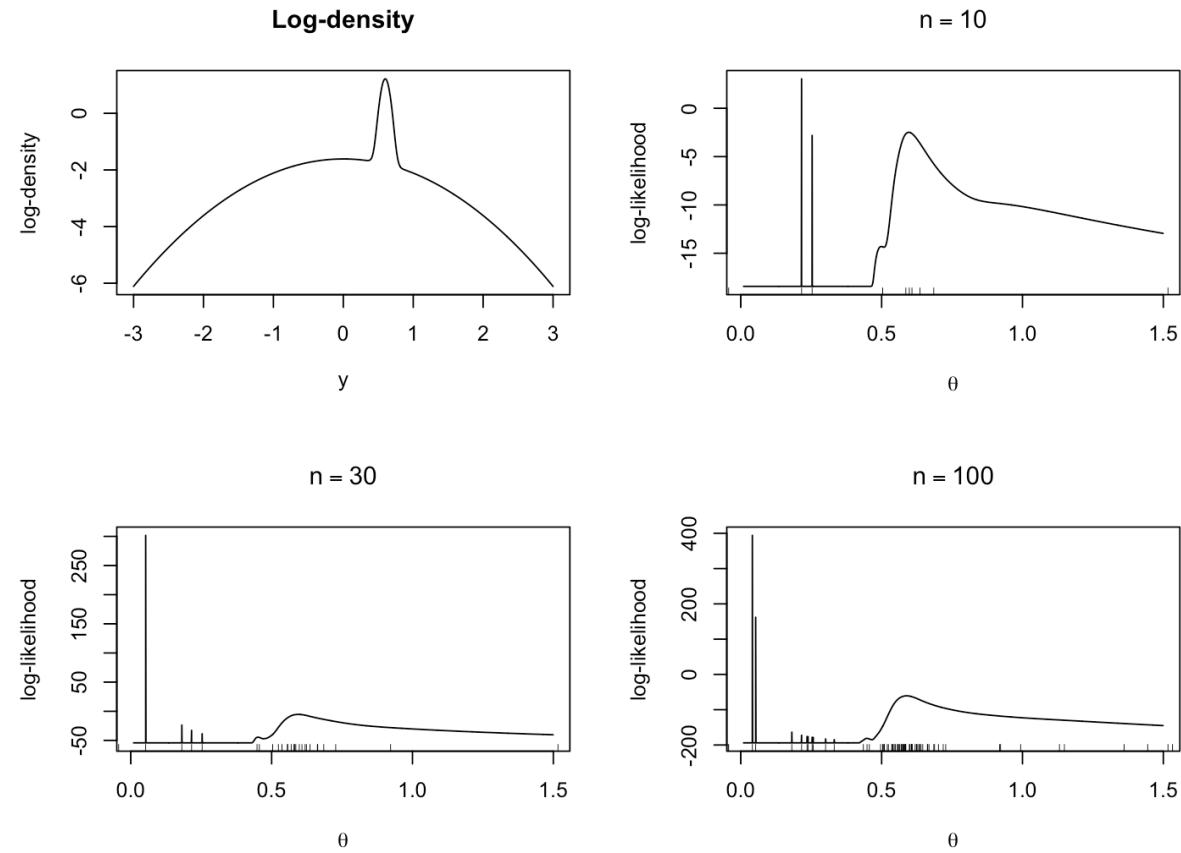
where  $\phi(x; \mu, \sigma^2)$  is the **normal density**. The density  $f(y; \theta)$  satisfies conditions **(A1)-(A4)**. Moreover  $\ell(\theta)$  is **differentiable**.

- However, the log-likelihood present **spykes** in correspondence of the observed data and the maximum likelihood concentrates around 0 rather than  $\theta_0$ , i.e  $\hat{\theta}_n$  is **inconsistent**.
- Indeed, there are **multiple roots** of the score equation  $\ell^*(\theta)$ , corresponding to the spykes. The “correct” solution close  $\theta_0$  is among them, but is not identified by the data even for large  $n$ .

Note this example does not contradict Wald inequality, because the spykes occur on a **set of measure zero**. Indeed for any fixed  $\theta$  the probability of observing  $Y_i = \theta$  is zero.

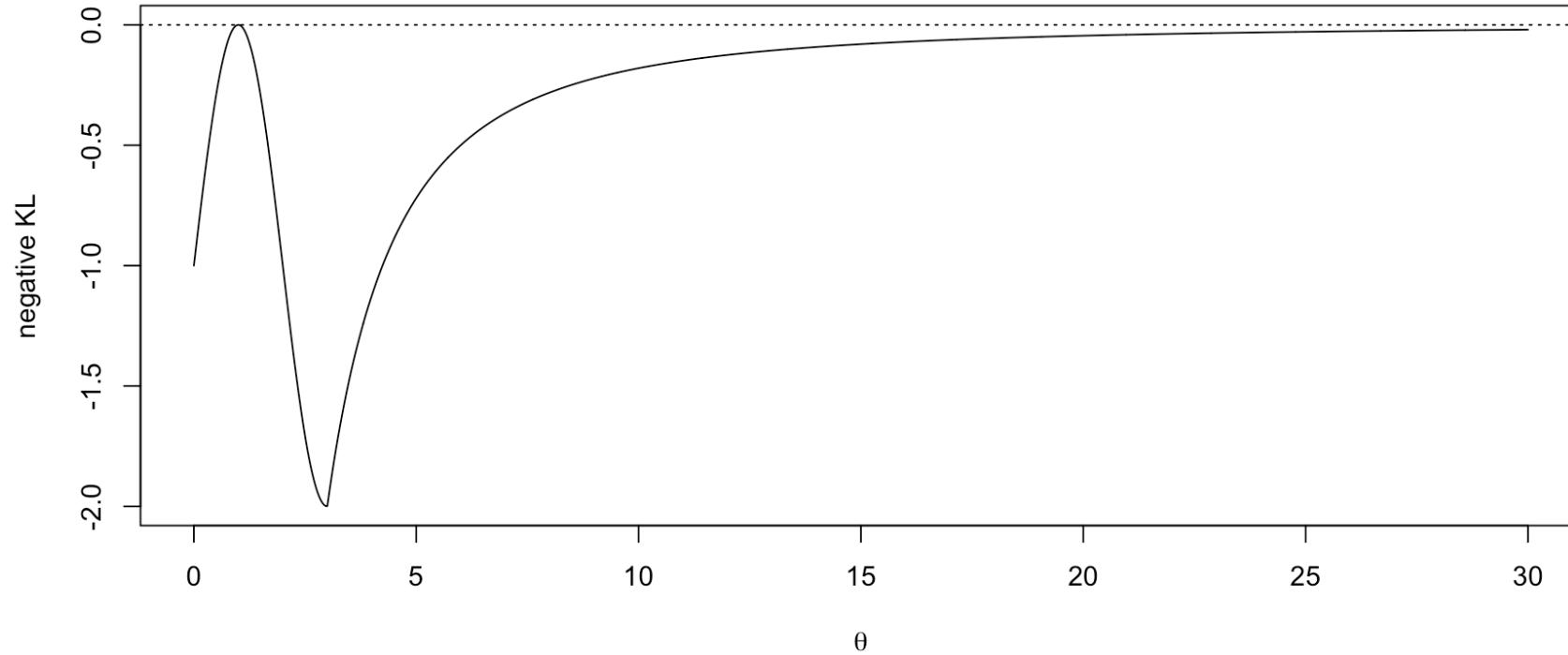
- This example is taken from a **blogpost** of Radford Neal.

# What could go wrong?



- These are **simulated data** when the true value is  $\theta_0 = 0.6$ . We plot the **log-density**  $\log f(y; \theta_0)$  and the **log-likelihood**  $\ell(\theta)$  for  $n = 10, 30, 100$ .

# What else could go wrong?



- The picture depict  $-\text{KL}(f(\cdot; \theta_0) \mid f(\cdot; \theta))$  in another **problematic** situation. The true value is  $\theta_0 = \pi/2$ , but the presence of the asymptote may cause  $\hat{\theta}_n$  to **diverge**.

# Consistency for the MLE

**Theorem (Lehmann and Casella (1998), Th. 5.1, Chap. 6)**

Under assumptions **(A1)-(A6)**, with probability tending to 1 as  $n \rightarrow \infty$ , the equation  $\ell^*(\theta) = \mathbf{0}$  has at least one root  $\hat{\theta}_n$ , and there exists a sequence of roots  $\hat{\theta}_n$  such that  $\hat{\theta}_n \xrightarrow{p} \theta_0$ .

- This standard result of the literature requires several regularity conditions and yet it delivers **less** than **what it seems**.<sup>1</sup>
- The claim is that a **clairvoyant statistician**, with knowledge of  $\theta_0$ , could choose a consistent sequence of roots. In reality, it may be impossible to choose the right solution.

Let us further require that, with probability tending to 1 as  $n \rightarrow \infty$ , there is a **unique solution** to the **score equation**. In that case  $\hat{\theta}_n$  is **consistent** and is also the **global maximizer** of  $\ell(\theta)$ .

1. See also van der Vaart (1998), Theorem 5.42, for slightly less stringent conditions and a careful discussion about the issue of multiple roots.

# Asymptotic normality of the MLE

**Theorem (Lehmann and Casella (1998), Th. 5.1, Chap. 6)**

Under assumptions **(A1)-(A6)**, suppose the maximum likelihood estimator  $\hat{\theta}_n$  exists and is **consistent** for the true value  $\theta_0$ . Then

$$\sqrt{n}(\hat{\theta}_n - \theta_0) \xrightarrow{d} N(0, i(\theta_0)^{-1}), \quad i(\theta) = - \int \left( \frac{\partial^2}{\partial \theta \partial \theta^T} \log f(y \mid \theta) \right) f(y \mid \theta) dy,$$

where  $ni(\theta)$  is the **Fisher information matrix**.

- Informally, we say that the maximum likelihood estimator is **asymptotically efficient** because, roughly speaking,  $\text{var}_\theta(\hat{\theta}_n) \approx I(\theta)^{-1}$ , the latter being the **Cramér-Rao lower bound**.
- Rigorously, the above theorem does not establish the convergence of  $\mathbb{E}_\theta(\hat{\theta}_n)$  and  $\text{var}_\theta(\hat{\theta}_n)$ , nor have we introduced an asymptotic version of the Cramér-Rao bound.
- Nevertheless, the statement that maximum likelihood estimators are **asymptotically efficient** is **correct**, as rigorously discussed in Chapter 8 of van der Vaart (1998).

# Observed vs Fisher information

- For the **practical** construction of **confidence intervals**, or simply to empirically assess the variance of an estimator, we need to consider a consistent estimator of  $I(\theta) = ni(\theta)$ .

The negative **Hessian matrix** of the log-likelihood is called **observed information** and equals:

$$\mathcal{I}(\theta) := -\frac{\partial^2}{\partial\theta\partial\theta^T}\ell(\theta; \mathbf{Y}).$$

If the second Bartlett identity holds, the **Fisher information** is  $I(\theta) = \mathbb{E}_\theta(\mathcal{I}(\theta))$ .

- There are two natural candidates for estimating  $I(\theta_0) = ni(\theta_0)$ , namely  $I(\hat{\theta}_n)$  and  $\mathcal{I}(\hat{\theta}_n)$ . These two quantities may coincide, but in general, this is not guaranteed.
- Following Fisher's original work, Efron and Hinkley (1978) suggested using  $\mathcal{I}(\hat{\theta}_n)$  because it approximates the **conditional variance** of  $\hat{\theta}_n$  given an appropriate **ancillary statistic**.
- Moreover,  $\mathcal{I}(\hat{\theta}_n)$  can be computed numerically via differentiation, whereas  $I(\hat{\theta}_n)$  requires analytical derivations.

# Consistency for M-estimators I

- We now discuss a broader and modern theory for **consistency**. Let  $M_n(\theta) = \sum_{i=1}^n m(\theta; Y_i)$  be an **M-estimator** and recall that the **maximum likelihood** is a special instance, with

$$M_n(\theta) = \sum_{i=1}^n \ell(\theta; Y_i),$$

- In order to simplify the subsequent exposition, it is convenient to consider

$$M_n(\theta) = \frac{1}{n} \sum_{i=1}^n [\ell(\theta; Y_i) - \ell(\theta_0; Y_i)] = \frac{1}{n} \sum_{i=1}^n \log \frac{f(Y_i; \theta)}{f(Y_i; \theta_0)},$$

where the red terms are ininfluent because the maximizer  $\hat{\theta}_n$  of  $M_n(\theta)$  is the same.

- Under conditions **(A1)-(A3)** the law of large numbers guarantees that  $M_n(\theta) \xrightarrow{p} M(\theta)$  **pointwise** for every  $\theta$ , where  $M(\theta) = -\text{KL}(f(\cdot; \theta_0) \mid f(\cdot; \theta))$ . However, this was **not enough**.
- On the other hand, we will see that  $\hat{\theta}_n$  does not need to maximize  $M_n(\theta)$ . Indeed, it is sufficient that is **nearly maximizes** it, in the sense that  $M_n(\hat{\theta}_n) \geq \sup_{\theta \in \Theta} M(\theta) - o_p(1)$ .

# Consistency for M-estimators II



## Theorem (van der Vaart (1998), Theorem 5.7)

Let  $\theta \subseteq \mathbb{R}^p$ ,  $M_n(\theta)$  be random functions and  $M(\theta)$  be a function of  $\theta$  such that for every  $\epsilon > 0$

$$\sup_{\theta \in \Theta} |M_n(\theta) - M(\theta)| \xrightarrow{p} 0, \quad (\text{Uniform convergence})$$

$$\sup_{\theta: \|\theta - \theta_0\|_2 \geq \epsilon} M(\theta) < M(\theta_0). \quad (\text{Strong identifiability})$$

Then any sequence of estimators  $\hat{\theta}_n$  with  $M_n(\hat{\theta}_n) \geq M_n(\theta_0) - o_p(1)$  **converges in probability** to  $\theta_0$  as  $n \rightarrow \infty$ .

- The **uniform convergence** assumption strengthens the pointwise convergence typically ensured by the law of large numbers, that is,  $m(\theta)$  should be **Glivenko-Cantelli**.
- This holds if  $\Theta$  is **compact**,  $m(\theta)$  is continuous and dominated by an integrable function.
- The strong identifiability condition, also called **well separability** ensures that only points close to  $\theta_0$  are close to the maximum value  $M(\theta_0)$ , strengthening Wald inequality.

# Consistency for Z-estimators

- The former theorem can be also expressed in terms of Z-estimators  $Q_n(\theta)$ , that is, on a set of **estimating equations**. An example is the score function

$$Q_n(\theta) = \frac{1}{n} \sum_{i=1}^n \ell^*(\theta; Y_i).$$

- We require  $\hat{\theta}_n$  to **nearly solve**  $Q_n(\theta) = \mathbf{0}$  and that  $\lim_{n \rightarrow \infty} Q_n(\theta_0) = \mathbf{0}$ . Intuitively, this comes from the **LLN** in **unbiased estimating equations**, in which case  $\mathbb{E}_{\theta_0}(Q_n(\theta_0)) = \mathbf{0}$ .

## Theorem (van der Vaart (1998), Theorem 5.9)

Let  $\theta \subseteq \mathbb{R}^p$ ,  $Q_n(\theta)$  be random vector-valued functions and  $Q(\theta)$  be a vector-valued function of  $\theta$  such that for every  $\epsilon > 0$

$$\sup_{\theta \in \Theta} \|Q_n(\theta) - Q(\theta)\|_2 \xrightarrow{P} 0, \quad \inf_{\theta: \|\theta - \theta_0\|_2 \geq \epsilon} \|Q(\theta)\|_2 > \|Q(\theta_0)\|_2 = 0.$$

Then any sequence of estimators  $\hat{\theta}_n$  such that  $Q_n(\hat{\theta}_n) = o_p(1)$  **converges in probability** to  $\theta_0$ .

# Asymptotic normality of M-estimators I

**Theorem (van der Vaart (1998), Theorem 5.21)**

Let  $\Theta \subseteq \mathbb{R}^p$  and  $M_n(\theta)$  be an M-estimator whose vector-valued **derivative** is  $Q_n(\theta)$ , so that

$$M_n(\theta) = \sum_{i=1}^n m(\theta; Y_i), \quad Q_n(\theta) = \sum_{i=1}^n q(\theta; Y_i).$$

Let  $\hat{\theta}_n$  be a sequence of **consistent** estimators such that  $Q_n(\hat{\theta}_n) = 0$ . Under assumptions **(A1)-(A2)** and further **mild regularity conditions**

$$\sqrt{n}(\hat{\theta}_n - \theta_0) \xrightarrow{d} N(0, v(\theta_0)^{-1}), \quad v(\theta) = h(\theta)j(\theta)^{-1}h(\theta),$$

where  $V(\theta) = nv(\theta)$  is the **Godambe information matrix** and

$$h(\theta) = \mathbb{E}_\theta \left( -\frac{\partial}{\partial \theta} q(\theta) \right), \quad j(\theta) = \mathbb{E}_\theta (q(\theta)q(\theta)^T).$$

## Asymptotic normality of M-estimators II

- The previous theorem is very powerful as it applies to the broad class of M-estimators. Moreover, its statement has been substantially **simplified** compared to van der Vaart (1998).
- The actual statement does not even require  $Y_1, \dots, Y_n$  to be independent or identically distributed.
- Moreover, the **regularity conditions** are much weaker than **(A1)-(A6)** and essentially ensure that the involved quantities are well-defined. **However**, note that  $\hat{\theta}_n$  must be **consistent**.
- An alternative proof relying on more **classical conditions**, e.g. based on bounding third derivatives as in **(A1)-(A6)**, is given in van der Vaart (1998), Theorem 5.41.

In regular problems, an M-estimator  $\hat{\theta}_n$  is **consistent** and **asymptotically normal**. Moreover, for  $n$  large enough, its variance is the inverse of the **Godambe information**. Informally,

$$\hat{\theta}_n \stackrel{d}{\sim} N(\theta_0, V(\theta_0)^{-1}).$$

# First-order bias-correction

- In a regular model with  $\Theta \subseteq \mathbb{R}^p$  and independent samples, the **bias** of the **maximum likelihood estimator** can be **expanded** as follows:

$$\text{bias}_\theta(\hat{\theta}_n) = \frac{b_1(\theta)}{n} + \frac{b_2(\theta)}{n^2} + \mathcal{O}(n^{-3}).$$

- The quantity  $\text{bias}_\theta(\hat{\theta}_n)$  is often unavailable, but the **first-order** term  $b_1(\theta)$  might be computable, e.g., in **exponential families** (Pace and Salvan (1997), Chap. 9).
- The **jackknife** is a popular strategy for removing the first-order bias. Alternatively, the **first-order bias-corrected** maximum likelihood estimator is obtained via plug-in as

$$\hat{\theta}_{\text{bc}} = \hat{\theta}_n - \frac{b_1(\hat{\theta}_n)}{n}.$$

- If  $\tilde{\theta}_n$  is an estimator of  $\theta$  with bias of order  $\mathcal{O}(n^{-2})$ , then

$$\text{var}_\theta(\tilde{\theta}_n) = \text{var}_\theta(\hat{\theta}_{\text{bc}}) + \Delta^2(\theta) + \mathcal{O}(n^{-3}),$$

where  $\Delta^2(\theta) \geq 0$ , with equality if and only if  $\hat{\theta}_{\text{bc}} = \tilde{\theta}_n$ . In this case,  $\hat{\theta}_{\text{bc}}$  is said to be **second-order efficient**; see Efron (1975).

## Bias reduction using Firth's correction I

- The jackknife and  $\hat{\theta}_{\text{bs}}$  are “corrective” rather than “preventive”. That is, the maximum likelihood  $\hat{\theta}$  is first calculated, then corrected. A practical requirement is the **existence** of  $\hat{\theta}$ .
- Motivated by this, Firth (1993) proposed a **modified score equation**. The idea is that the **bias** in  $\hat{\theta}$  can be **reduced** by introducing a **small bias** into the **score function**.

In a regular model with  $\Theta \subseteq \mathbb{R}^p$ , let  $\ell^*(\theta)$  be the score function,  $I(\theta)$  the Fisher information matrix and  $b(\theta)$  the **bias** of the maximum likelihood  $\hat{\theta}_n$ . Firth (1993) estimating equation is

$$Q(\theta) = \ell^*(\theta) + A(\theta),$$

where  $A(\theta) = A(\theta; \mathbf{Y})$  is any vector such that

$$\mathbb{E}_\theta(A(\theta)) = -I(\theta) \frac{b_1(\theta)}{n} + \mathcal{O}(n^{-1/2}).$$

Clearly, a natural candidate for  $A(\theta; \mathbf{Y})$  is indeed  $A(\theta) = -I(\theta)b_1(\theta)/n$ .

## Bias reduction using Firth's correction II

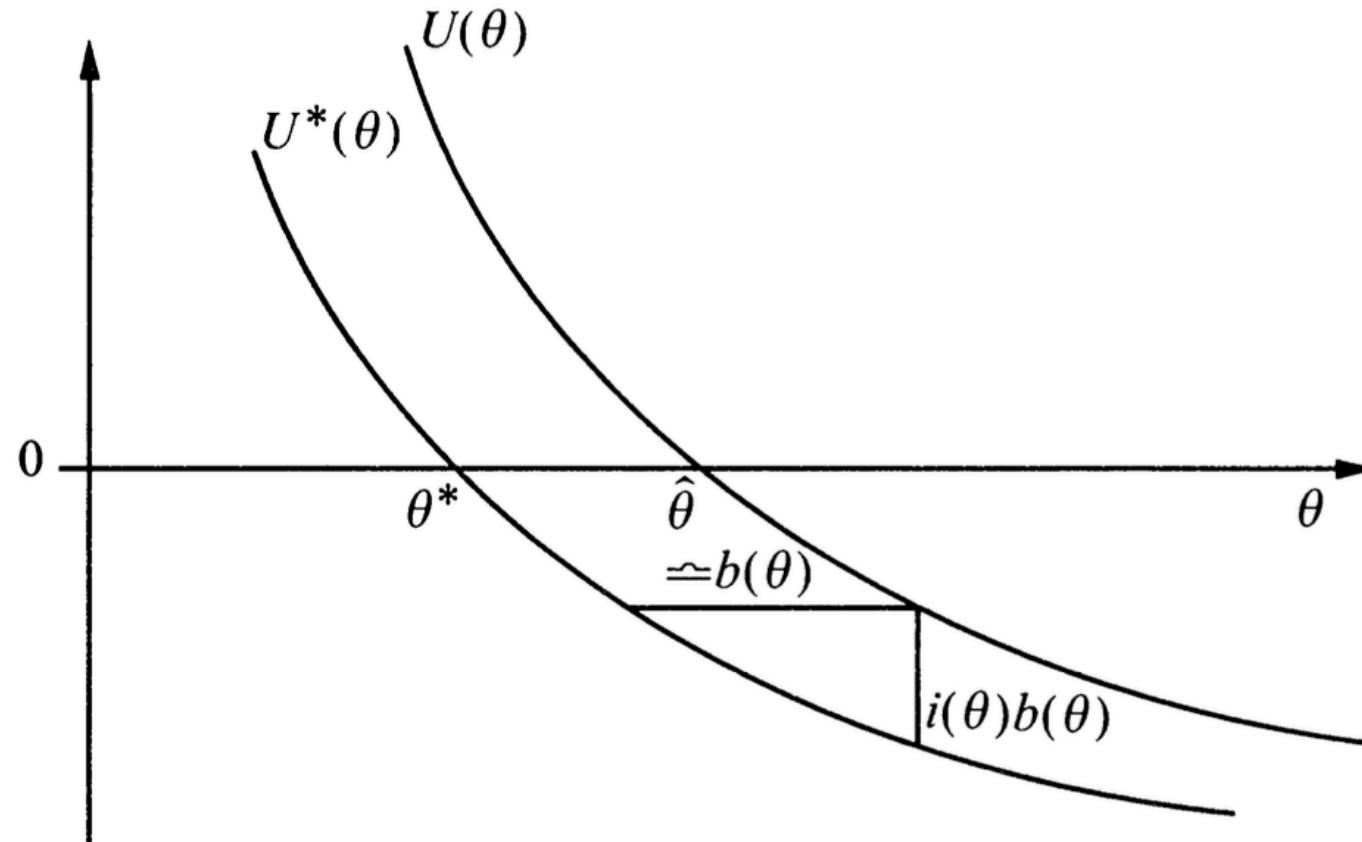


Fig. 1. Modification of the unbiased score function.

## Bias reduction using Firth's correction III



Let  $Y_1, \dots, Y_n$  be an iid sample from a Poisson with parameter  $\lambda$  and consider the reparametrization  $\psi = 1/\lambda$ . The score and the Fisher information are

$$\ell^*(\psi) = \frac{n}{\psi^2} - \frac{\bar{y}}{n\psi}, \quad I(\psi) = \frac{n}{\psi^3}.$$

The **maximum likelihood** is  $\hat{\psi} = 1/\bar{y}$ , with **first-order bias**  $b_1(\psi)/n = \psi^2/n$ . Thus

$$Q(\psi) = \ell^*(\psi) - I(\psi)b_1(\psi)/n \implies \hat{\psi}_{\text{Firth}} = \frac{1}{\bar{y} + 1/n}.$$

Let  $Y_1, \dots, Y_n$  be an iid sample from a Bernoulli with parameter  $p$  and consider the reparametrization  $\beta = \log p/(1 - p)$ . Application of Firth (1993) method gives

$$\hat{\beta}_{\text{Firth}} = \log \left( \frac{n_1 + 1/2}{n - n_1 + 1/2} \right), \quad \text{whereas} \quad \hat{\beta} = \log \left( \frac{n_1}{n - n_1} \right),$$

which is a well-known **bias-reducing** correction of the empirical logit.

# Robustness

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# Robustness: preliminaries

- Thus far, we have assumed that  $\mathcal{F} = \{f(\cdot; \theta) : \theta \in \Theta\}$  is **correctly specified**, that is, there exists a  $\theta_0$  that generates the data  $(Y_1, \dots, Y_n) \sim f(\cdot; \theta_0)$  and  $f(\cdot; \theta_0) \in \mathcal{F}$ .
- Under this **assumption**, we have derived estimators that are **optimal** in some sense. However, if the underlying model is not correct, everything breaks down.
- The term “**robustness**” is intentionally vague, but let us say that any statistical procedure:
  - a. Should have nearly optimal efficiency if the model is correctly specified
  - b. Small deviations from the model assumptions should impact the model only slightly
  - c. Somewhat larger deviations from the model should not cause a catastrophe
- We also distinguish among two kinds of robustness:
  - i. robustness with respect to **contamination** of the data (i.e.  $y_i = 10^{32}$  is an **outlier**)
  - ii. robustness with respect to **model misspecification**, that is, we specify a class of models  $\mathcal{F}$  but in reality  $(Y_1, \dots, Y_n) \sim f_0(\cdot)$  and  $f_0(\cdot) \notin \mathcal{F}$ .
- Case i. is sometimes called **resistence** and relies on the notion of **influence functions**. For instance, the median is resistant, the mean is not.

## Example: Huber estimators I

- Recall that a Huber estimator  $\hat{\theta}_n$  for the **mean**  $\theta_0$  is defined as the solution of

$$Q(\theta) = \sum_{i=1}^n q(Y_i - \theta) = 0, \quad q(y) = \begin{cases} -k & \text{if } y \leq -k \\ y & \text{if } |y| \leq k \\ k & \text{if } y \geq k \end{cases}.$$

- We assume  $Y_i \stackrel{\text{iid}}{\sim} f_0$  where  $f_0(y)$  is a **continuous** and **symmetric** density around  $\theta_0$ . Hence, there exists a density  $\tilde{f}_0$  symmetric around 0 such that  $\tilde{f}_0(y - \theta_0) = f_0(y)$ .
- First of all, provided  $f_0$  is symmetric, the estimating equation  $Q(\theta)$  is **unbiased**

$$\mathbb{E}_0\{Q(\theta)\} = n\mathbb{E}_0\{q(Y_1 - \theta_0)\} = 0.$$

Broadly speaking, this means that Huber estimator  $\hat{\theta}_n$  is **consistent** for  $\theta_0$ . Moreover

$$\sqrt{n}(\hat{\theta}_n - \theta_0) \xrightarrow{d} \mathcal{N}(0, v^{-1}(\theta_0)), \quad v(\theta)^{-1} = j(\theta)/h(\theta)^2.$$

- Huber estimator is **robust** but **less efficient** than the maximum likelihood. For instance, if  $f_0$  is a Gaussian with mean  $\theta_0$  and variance  $\sigma^2$ , then  $\bar{Y}$  has asymptotic variance  $\sigma^2$  and  $\sigma^2 \leq v(\theta_0)^{-1}$ .

## Example: Huber estimators II

- Let  $Z \sim \tilde{f}_0$ , and note such density **does not depend** on  $\theta_0$ . After some calculations, we find

$$j(\theta) = \mathbb{E}_0\{q(Y_1 - \theta)^2\} = \mathbb{E}\{Z^2 I(|Z| < k)\} + 2k^2 \mathbb{P}(Z > k),$$

and

$$h(\theta) = \mathbb{E}_0\left(-\frac{\partial}{\partial \theta}q(Y_1 - \theta)\right) = \mathbb{P}(|Z| \leq k).$$

- Remarkably, the asymptotic variance does not depend on  $\theta_0$  therefore its relative efficiency compared to the **normal model** is also **constant** over  $\theta_0$ .
- Below we show the **asymptotic relative efficiency** (ARE) of the Huber estimator compared to  $\bar{Y}$  for different values of  $k$ , assuming  $f_0$  is Gaussian with  $\sigma^2 = 1$  and arbitrary  $\theta_0$

$k$	0 (Median)	0.5	1	1.5	2
$ARE = \sigma^2/v(\theta_0)^{-1} = v(\theta_0)$	0.637	0.792	0.903	0.964	0.99

- The ARE does not depend on  $\sigma^2$  if we use  $k = \sigma\tilde{k}$ . This is the default of the **huber** function of the **MASS** R package with  $\tilde{k} = 1.5$ , where  $\sigma$  is robustly estimated using the MAD.

## Example: omitted variable in linear regression

- Let  $y_1, \dots, y_n$  be realizations from the model  $Y_i = \beta_1 + \beta_2 x_i + \beta_3 z_i + \epsilon_i$  for  $i = 1, \dots, n$ , where  $x_i$  and  $z_i$  are linearly independent **covariates** and  $\epsilon_i \stackrel{\text{iid}}{\sim} N(0, \sigma^2)$  is the **error term**.
- If we do not include  $z_i$  in our model, the class  $\mathcal{F}$  is **misspecified**. The **maximum likelihood** estimate of  $\beta_2$  under the misspecified  $\mathcal{F}$  is:

$$\hat{\beta}_2 = \frac{1}{\sum_{j=1}^n (x_j - \bar{x})^2} \sum_{i=1}^n (x_i - \bar{x})(y_i - \bar{y}).$$

- Thus, the estimator  $\hat{\beta}_2$ , **under the true model**  $f_0$ , is distributed as a Gaussian with mean

$$\mathbb{E}_0(\hat{\beta}_2) = \beta_2 + \beta_3 \frac{1}{\sum_{j=1}^n (x_j - \bar{x})^2} \sum_{i=1}^n (x_i - \bar{x})(z_i - \bar{z}),$$

and variance  $\sigma^2 / \sum_{j=1}^n (x_j - \bar{x})^2$ .

- In other words,  $\hat{\beta}_2$  is **biased** and **inconsistent** unless  $x$  and  $z$  are **uncorrelated** or, obviously, if the model is correctly specified, that is if  $\beta_3 = 0$ .

# Maximum likelihood under a misspecified model



- Let  $f_0(\cdot)$  denote the true probability model, and let  $\mathcal{F} = \{f(\cdot; \theta) : \theta \in \Theta\}$  be an **incorrectly specified** statistical model, such that  $f_0 \notin \mathcal{F}$ .
- Suppose  $Y_1, \dots, Y_n$  are **iid** under  $f_0$ , and define the log-likelihood as  $\ell(\theta) = \sum_{i=1}^n \log f(y_i; \theta)$ .
- Let  $\mathbb{E}_0(\cdot)$  denote expectation under the true model  $f_0$ . If the problem is **sufficiently regular**, the **maximum likelihood** estimator  $\hat{\theta}_n$  **converges** in probability to some value  $\theta_0$  such that

$$\mathbb{E}_0\{\ell(\theta)\} < \mathbb{E}_0\{\ell(\theta_0)\}, \quad \text{for all } \theta \in \Theta, \quad \theta \neq \theta_0,$$

as shown in Huber (1967). That is,  $\hat{\theta}_n$  converges to a value satisfying **Wald's inequality**.

- An alternative formulation of this result is the following:

$$\text{KL}(f(\cdot; \theta_0) \mid f_0) < \text{KL}(f(\cdot; \theta) \mid f_0), \quad \text{for all } \theta \in \Theta, \quad \theta \neq \theta_0.$$

In other words, the **maximum likelihood** converges to  $\theta_0$ , which represents the parameter value that makes  $f(\cdot; \theta)$  **as close as possible** to the true  $f_0$ , albeit  $\text{KL}(f(\cdot; \theta_0) \mid f_0) > 0$ .

The maximum likelihood makes our predictions “the best they can be” given the chosen model.

## Example: misspecified exponential model I

- Let  $\mathcal{F} = \{\mu^{-1}e^{-y/\mu} : \mu \in \mathbb{R}^+\}$  and suppose the data  $Y_1, \dots, Y_n$  are iid from an **exponential** with mean  $\mu_0$ . Then  $\bar{Y}$  is a **consistent** and **efficient** estimator for  $\mu_0$ .
- However,  $\bar{Y}$  is a **robust** estimator, in the sense that if instead  $Y_i \stackrel{\text{iid}}{\sim} f_0(\cdot)$  with  $f_0(\cdot) \notin \mathcal{F}$ , then  $\bar{Y}$  remains **consistent** for the mean  $\mu_0$  under minimal assumptions on  $f_0$  (i.e.  $\mu_0$  must exist).
- On the other hand, the central limit theorem shows that the asymptotic distribution is

$$\sqrt{n}(\bar{Y} - \mu) \xrightarrow{d} N(0, \sigma^2).$$

where  $\sigma^2$  is the variance of  $Y_i$  under  $f_0(\cdot)$ , provided  $\sigma^2 < \infty$ .

- Thus, **confidence intervals** are robust if  $\sigma^2$  is estimated in a **robust** way, e.g. using the **sample variance**, but **not** if we use the usual  $\hat{\sigma} = i(\hat{\mu})^{-1} = \bar{Y}^2$  implied by the exponential specification.
- This example can be read under the lenses of estimating equations. The score function

$$\ell^*(\mu) = \sum_{i=1}^n (Y_i - \mu)$$

is an **unbiased estimating equation** under  $f_0(\cdot)$  for a broad class of models beyond  $\mathcal{F}$ .

# Robustness and unbiased estimating equations

An essential requirement is that estimands must have the **same interpretation** under all the potential models. In the former exponential example  $\mu_0$  represents the mean of  $f_0$ .

Formally, this means we can write the parameter  $\theta_0$  as a **functional of interest**  $T(\cdot)$  of  $f_0$ :

$$\theta_0 = T(f_0).$$

For example, we might have  $\theta_0 = \int_{\mathcal{Y}} \mathbf{y} f_0(\mathbf{y}) \nu(d\mathbf{y})$ , that is, the mean of  $f_0$ .

On the other hand, for instance the parameters  $\alpha$  and  $\beta$  of a Gamma distribution are not robust to interpretation, because they are meaningless for models other than the Gamma.

- More broadly, if the **score function**  $\ell^*(\theta) = \sum_{i=1}^n \ell^*(\theta; Y_i)$  is **unbiased** under  $f_0(\cdot)$ , the estimator is **consistent** under mild conditions.
- Moreover, even if the maximum likelihood  $\hat{\theta}_n$  is consistent, the **asymptotic variance** is not anymore the one induced by the Fisher information: **adjustments** are needed.

# Misspecified likelihoods are M-estimators

- The **theory of M- and Z-estimators** can be directly applied to investigate the asymptotic behavior the maximum likelihood estimator **under model misspecification**.
- If  $\mathcal{F}$  is misspecified then the maximizer of the log-likelihood  $\ell(\theta) = \sum_{i=1}^n \log f(y_i; \theta)$  should be regarded as an **M-estimator** while the score function  $\ell^*(\theta)$  is a **Z-estimator**.
- **Consistency** and **asymptotic normality** of the maximum likelihood estimator  $\hat{\theta}_n$  for misspecified models hold under the assumptions in van der Vaart (1998), Theorems 5.7, 5.9, 5.21.
- Note that under misspecification, **Bartlett identity** does not hold anymore. However, under the assumptions of Theorem 5.21 of van der Vaart (1998), the maximum likelihood is such that

$$\sqrt{n}(\hat{\theta}_n - \theta_0) \xrightarrow{d} N(0, v(\theta_0)^{-1}), \quad v(\theta) = h(\theta)j(\theta)^{-1}h(\theta),$$

where  $V(\theta) = nv(\theta)$  is the **Godambe information matrix** and

$$h(\theta) = \mathbb{E}_0 \left( -\frac{\partial}{\partial \theta} \ell^*(\theta) \right), \quad j(\theta) = \mathbb{E}_0 \left( \ell^*(\theta) \ell^*(\theta)^T \right).$$

where expectations are taken over  $f_0$  and not the misspecified  $f(\cdot; \theta)$ , so that  $h(\theta) \neq j(\theta)$ .

# Sandwich estimators

- In order to compute **confidence intervals** and test hypotheses, we need to estimate the **asymptotic variance** of the maximum likelihood estimator  $\hat{\theta}_n$  under model misspecification.
- Informally, recall that for  $n$  large enough and under regularity conditions, we have

$$\text{var}_0(\hat{\theta}_n) \approx V(\theta_0)^{-1}.$$

- If the model is correctly specified, then typical **estimators** of the **variance** of  $\hat{\theta}_n$  are  $\mathcal{I}(\hat{\theta}_n)$  and  $I(\hat{\theta}_n)$ . Unfortunately, we cannot use  $V(\hat{\theta}_n)$  because it depends on  $f_0$ , which is unknown!

The **sandwich estimator** is a popular choice for estimating the **asymptotic variance** of  $\hat{\theta}_n$  under model misspecification. The sandwich estimator for  $V(\theta)$  is

$$\hat{V}(\hat{\theta}_n) = \mathcal{I}(\hat{\theta}_n) \left( \sum_{i=1}^n \ell^*(\hat{\theta}_n; y_i) \ell^*(\hat{\theta}_n; y_i)^T \right)^{-1} \mathcal{I}(\hat{\theta}_n),$$

recalling that  $\mathcal{I}(\theta) = -\partial/\partial\theta \ell^*(\theta; \mathbf{y})$  is the **observed information** matrix. If the model is correctly specified then  $\hat{V}(\hat{\theta}_n) = \mathcal{I}(\hat{\theta}_n) + o_p(1)$ .

## Example: misspecified exponential model II

- Let us consider again  $Y_i \stackrel{\text{iid}}{\sim} f_0(\cdot)$  with  $f_0(\cdot) \notin \mathcal{F}$ , where  $\mathcal{F}$  is a **misspecified exponential** model.
- The score function is **unbiased** for  $\mu$  under  $f_0$ . The score and the observed observation matrix are, respectively

$$\ell^*(\mu) = \sum_{i=1}^n \left( -\frac{1}{\mu} + \frac{y_i}{\mu^2} \right) = -\frac{n}{\mu} + \frac{n\bar{y}}{\mu^2}, \quad \mathcal{I}(\mu) = -\frac{\partial}{\partial \mu} \ell^*(\mu) = -\frac{n}{\mu^2} + \frac{2n\bar{y}}{\mu^3}.$$

- The inverse of the observed information matrix, evaluated at  $\hat{\mu} = \bar{y}$ , is an estimate of the asymptotic variance **assuming** the model is **correctly specified**:

$$\text{var}_\theta(\hat{\mu}) \approx \mathcal{I}(\hat{\mu})^{-1} = \frac{\bar{y}^2}{n}.$$

- If the model is **misspecified**, then we can use the **sandwich estimator**:

$$\text{var}_0(\hat{\mu}) \approx \mathcal{I}(\hat{\mu})^{-2} \sum_{i=1}^n \ell^*(\hat{\mu}; y_i)^2 = \frac{\bar{y}^4}{n^2} \sum_{i=1}^n \frac{1}{\bar{y}^4} (y_i - \bar{y})^2 = \frac{1}{n} \left( \frac{1}{n} \sum_{i=1}^n (y_i - \bar{y})^2 \right),$$

which is the  $n^{-1}$  times the usual method of moments estimator for the variance.

## Example: correlated observations

- Let  $y_1, \dots, y_n$  be realizations of **dependent** random variables from the density  $f(\mathbf{y}; \theta)$  with  $\Theta \subseteq \mathbb{R}^p$ . Moreover, let  $f(y_i; \theta)$  be the **marginal density** of  $y_i$ .
- If the data are dependent, the log-likelihood  $\ell_c(\theta) = \sum_{i=1}^n \log f(y_i; \theta)$  is **misspecified**, because it assumes independence. Indeed,  $\ell_c(\theta)$  is an instance of **composite likelihood**.
- Nonetheless, under mild regularity conditions, the **misspecified score function** is **unbiased** under the true model  $f(\mathbf{y}; \theta)$ , thanks to the **linearity** of the expectation operator:

$$\ell_c^*(\theta) = \sum_{i=1}^n \frac{\partial}{\partial \theta} \log f(y_i; \theta) \implies \mathbb{E}_0\{\ell_c^*(\theta)\} = \mathbf{0}.$$

- Unbiasedness of  $\ell_c^*(\theta)$  does not guarantee consistency nor asymptotic normality, but both are recovered under mild assumptions on the dependence structure, relying on **ergodic** theorems or **martingales**. Remarkably, a Godambe-like asymptotic variance is obtained at the limit<sup>1</sup>.

- Refer to Section 7.2.3 in Davison (2003) for a discussion on general estimating equations with dependent data and sufficient conditions for consistency and asymptotic normality.

# Example: the probability of observing a zero

- Let  $Y_1, \dots, Y_n$  be an **iid** sample from a **discrete** distribution with pdf  $f_0$ . We are interested in estimating the following functional

$$\psi_0 = \mathbb{P}(Y_i = 0) = f_0(0),$$

- Under a Poisson model  $\mathcal{F}$  with mean  $\lambda$ ,  $\psi$  is reparametrization:  $\psi = e^{-\lambda}$ . Two estimators are:

$$\hat{\psi}_{\text{ML}} = e^{-\bar{y}} \quad (\text{maximum likelihood}), \quad \hat{\psi} = \left( \frac{n-1}{n} \right)^{n\bar{y}}, \quad (\text{UMVU}).$$

- Both estimators are **not robust** for  $\psi_0$  under model misspecification, unless under  $\psi_0$  we have  $\psi_0 = e^{-\mu_0}$ . Indeed, the score equation is **biased**, being equal to

$$\ell^*(\psi) = \frac{n}{\psi} \left( 1 - \frac{\bar{y}}{-\log \psi} \right) \implies \mathbb{E}_0\{\ell^*(\psi_0)\} = \frac{n}{\psi_0} \left( 1 - \frac{\mu_0}{-\log \psi_0} \right).$$

- A **robust alternative** is given by the **empirical proportion** of zero:

$$\hat{\psi}_{\text{MM}} = \frac{1}{n} \sum_{i=1}^n \mathbb{I}(y_i = 0).$$

## Example: linear models with misspecified variance I

- Let  $Y_i = \mathbf{x}_i^T \beta + \epsilon_i$ , where the **errors**  $\epsilon_i$  are random variables distributed according to  $f_0$ , whereas  $\mathbf{x}_i$  are **known covariates** and  $\beta \in \mathbb{R}^p$  is a **parameter** of interest.
- If we further assume that  $\epsilon_i \stackrel{\text{iid}}{\sim} N(0, \sigma^2)$ , then the **maximum likelihood estimator**  $\hat{\beta} = (\mathbf{X}^T \mathbf{X})^{-1} \mathbf{X}^T \mathbf{y}$  is the **least squares estimator**. Moreover, the **score function** is

$$\ell^*(\beta) = \frac{1}{\sigma^2} \sum_{i=1}^n \mathbf{x}_i (y_i - \mathbf{x}_i^T \beta) = \frac{1}{\sigma^2} \mathbf{X}^T (\mathbf{y} - \mathbf{X} \beta),$$

whereas the Fisher/observed **information matrix** is

$$I(\beta) = \mathcal{I}(\beta) = \frac{1}{\sigma^2} \sum_{i=1}^n \mathbf{x}_i \mathbf{x}_i^T = \frac{1}{\sigma^2} \mathbf{X}^T \mathbf{X} \implies \text{var}_\beta(\hat{\beta}) = \sigma^2 (\mathbf{X}^T \mathbf{X})^{-1}.$$

- Provided that  $\mathbb{E}_0(Y_i) = \mathbf{x}_i^T \beta$ , that is, as long as the **linearity assumption** holds under  $f_0$ , then the OLS estimator is **robust** and **unbiased** even when  $\epsilon_i$  are not Gaussian or **not iid**.
- If the iid Gaussian model is **misspecified**, e.g. because  $\text{var}_0(\epsilon_i) = \sigma_i^2$  (heteroskedasticity), then the OLS estimator is still a good choice but its variance should be adjusted.

## Example: linear models with misspecified variance II



- The **variance** of  $\hat{\beta}$  under a general model  $f_0$  in which the random vector  $(\epsilon_1, \dots, \epsilon_n)$  has zero mean and  $\text{var}_0(\epsilon) = \Sigma$ , is explicitly available

$$\text{var}_0(\hat{\beta}) = (\mathbf{X}^T \mathbf{X})^{-1} \mathbf{X}^T \Sigma \mathbf{X} (\mathbf{X}^T \mathbf{X})^{-1}.$$

This coincides with the inverse **Godambe information** of the **estimating equation**  $\ell^*(\beta)$ :

$$\text{var}_0(\hat{\beta}) = I(\beta)^{-1} \mathbb{E}_0 \left\{ \frac{1}{\sigma^4} \mathbf{X}^T (\mathbf{Y} - \mathbf{X}\beta) (\mathbf{Y} - \mathbf{X}\beta) \mathbf{X} \right\} I(\beta)^{-1}$$

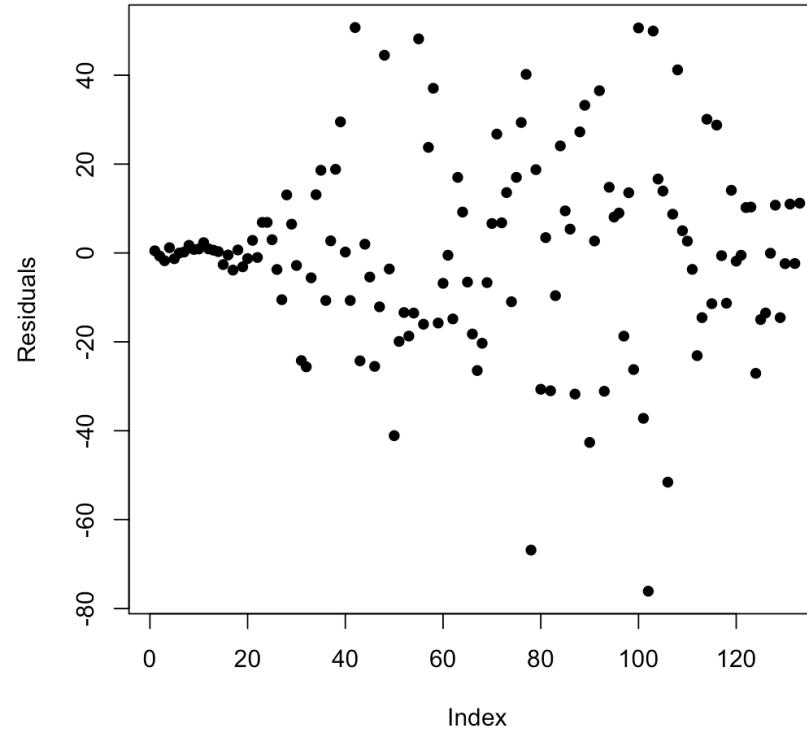
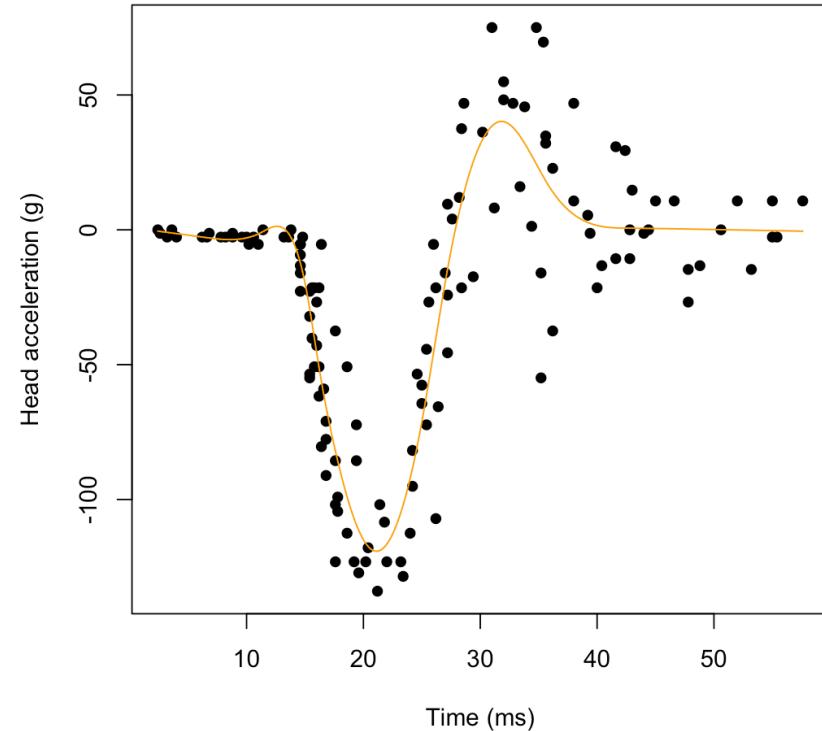
- In practice, the matrix  $\Sigma$  is unknown. Thus, we rely on the **sandwich estimator**:

$$\begin{aligned} \hat{V}(\hat{\beta})^{-1} &= I(\hat{\beta})^{-1} \left\{ \frac{1}{\sigma^4} \sum_{i=1}^n \mathbf{x}_i (y_i - \mathbf{x}_i^T \hat{\beta}) (y_i - \mathbf{x}_i^T \hat{\beta}) \mathbf{x}_i^T \right\} I(\hat{\beta})^{-1} \\ &= (\mathbf{X}^T \mathbf{X})^{-1} \mathbf{X}^T \mathbf{R}^2 \mathbf{X} (\mathbf{X}^T \mathbf{X})^{-1}, \end{aligned}$$

where  $\mathbf{R}^2 = \text{diag}(r_1^2, \dots, r_n^2)$  and  $r_i = y_i - \mathbf{x}_i^T \hat{\beta}$  are the **residuals**.

- This **robust** estimator is known as the **White's correction** after White (1980), who also proved it is consistent if  $\Sigma = \text{diag}(\sigma_1^2, \dots, \sigma_n^2)$  and under technical conditions on  $\mathbf{X}$ .

## Example: linear models with misspecified variance III



- In this example, we fit a linear model (in the parameters!) so that  $\mathbf{Y} = \mathbf{X}\beta + \epsilon$  using **least squares**, with  $p = 12$ . The residuals clearly show **heteroskedasticity**.

## Example: linear models with misspecified variance IV

- The **standard** estimator is  $\hat{\sigma}^2(\mathbf{X}^T \mathbf{X})^{-1}$ . Below we show the first  $4 \times 4$  entries

	$\beta_1$	$\beta_2$	$\beta_3$	$\beta_4$
$\beta_1$	<b>113.11</b>	-47.95	17.70	-12.81
$\beta_2$	-47.95	<b>170.53</b>	-73.32	54.10
$\beta_3$	17.70	-73.32	<b>87.43</b>	-69.35
$\beta_4$	-12.81	54.10	-69.35	<b>217.58</b>

- The **sandwich (White)** estimator is instead  $(\mathbf{X}^T \mathbf{X})^{-1} \mathbf{X}^T \mathbf{R}^2 \mathbf{X} (\mathbf{X}^T \mathbf{X})^{-1}$ :

	$\beta_1$	$\beta_2$	$\beta_3$	$\beta_4$
$\beta_1$	<b>2.82</b>	-9.73	9.45	-8.80
$\beta_2$	-9.73	<b>42.79</b>	-43.71	40.57
$\beta_3$	9.45	-43.71	<b>72.01</b>	-67.22
$\beta_4$	-8.80	40.57	-67.22	<b>288.72</b>

- We used the **sandwich** R package, that implements several sandwich variants.  
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# References and study material

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# Main references

- Davison (2003)
  - **Chapter 4** (*Likelihood*)
  - **Chapter 7** (*Estimation and Hypothesis Testing*)
  - **Chapter 8** (*Linear Regression Model*)
- Casella and Berger (2002)
  - **Chapter 7** (*Point estimation*)
  - **Chapter 10** (*Asymptotic evaluations*)
- Pace and Salvan (1997)
  - **Chapter 2** (*Data and model reduction*)
  - **Chapter 3** (*Survey of basic concepts and techniques*)

## Additional references

- Lehmann and Casella (1998)
  - **Chapter 1** (*Preparations*)
  - **Chapter 2** (*Unbiasedness*)
  - **Chapter 4** (*Average risk optimality*)
  - **Chapter 5** (*Minimax and admissibility*)
  - **Chapter 5** (*Minimax and admissibility*)
  - **Chapter 6** (*Asymptotic optimality*)
- Robert (1994)
  - **Chapter 2** (*Decision-theoretic foundations*)
- van der Vaart (1998)
  - **Chapter 4** (*Moment estimators*)
  - **Chapter 5** (*M- and Z-Estimators*)

# References

Agresti, A. (2015), *Foundations of Linear and Generalized Linear Models*, Wiley.

Casella, G., and Berger, R. L. (2002), *Statistical Inference*, Duxbury.

Davison, A. C. (2003), *Statistical Models*, Cambridge University Press.

The Annals of Statistics, 7, 269–292.

Efron, B. (1975), "Defining the curvature of a statistical problem (with applications to second order efficiency)," *The Annals of Statistics*, 3, 1189–1242.

Efron, B. (2010), *Large Scale Inference: Empirical Bayes Methods for Estimation, Testing, and Prediction*, Institute of mathematical statistics monographs, Leiden: Cambridge University Press.

Efron, B., and Hastie, T. (2016), *Computer Age Statistical Inference*, Cambridge University Press.

Efron, B., and Hinkley, D. V. (1978), "Assessing the accuracy of the maximum likelihood estimator: Observed vs fisher information," *Biometrika*, 65, 457–482.

Efron, B., and Morris, C. (1975), "Data analysis using Stein's estimator and its generalizations," *Journal of the American Statistical Association*, 70, 311–319.

Firth, D. (1993), "Bias reduction of maximum likelihood estimates," *Biometrika*, 80, 27–38.

Godambe, V. P. (1960), "An optimum property of regular maximum likelihood estimation," *Annals of Mathematical Statistics*, 31, 1208–1211.

Huber, P. J. (1967), "The behavior of maximum likelihood estimates under nonstandard conditions," in *Proceedings of the fifth berkeley symposium on mathematical statistics and probability*.