

Statistics: Homework 2

6.3 Given $\hat{\theta} = 2\bar{X}_n$ and $X_1, \dots, X_n \sim \text{Uniform}(0, \theta)$,

$$\begin{aligned}\text{bias}(\hat{\theta}) &= \mathbb{E}(2\bar{X}_n) - \theta \\ &= 2n^{-1} \mathbb{E} \left(\sum_{i=1}^n X_i \right) - \theta \\ &= 2n^{-1} \sum_{i=1}^n \mathbb{E}(X_i) - \theta \\ &= 2n^{-1} \frac{n\theta}{2} - \theta = 0 \\ \text{se}(\hat{\theta})^2 &= \mathbb{V}(2\bar{X}_n) \\ &= 4\mathbb{V}(\bar{X}_n) \\ &= 4n^{-2} \mathbb{V} \left(\sum_{i=1}^n X_i \right) \\ &= 4n^{-2} \sum_{i=1}^n \mathbb{V}(X_i) \\ &= 4n^{-2} \frac{n\theta^2}{12} = \frac{\theta^2}{3n} \\ \text{MSE}(\hat{\theta}) &= \text{bias}(\hat{\theta})^2 + \text{se}(\hat{\theta})^2 = \frac{\theta^2}{3n}\end{aligned}$$

7.2 For $X_1, \dots, X_n \sim \text{Bernoulli}(p)$ plug-in estimator for p is

$$\hat{p} = \frac{1}{n} \sum_{i=1}^n X_i$$

and the estimated standard error is given by

$$\hat{\text{se}}_p = \sqrt{\mathbb{V}(\hat{p})} = \sqrt{\frac{\hat{p}(1 - \hat{p})}{n}}$$

As the X_i 's are iid, by Central Limit Theorem, \hat{p} is asymptotically normal with mean p and variance $\hat{\text{se}}_p^2$. Thus an approximate 90% confidence interval for p is $(\hat{p} - 1.645\hat{\text{se}}_p, \hat{p} + 1.645\hat{\text{se}}_p)$.

For $X_1, \dots, X_n \sim \text{Bernoulli}(p)$ and $Y_1, \dots, Y_m \sim \text{Bernoulli}(q)$ plug-in estimator for $p - q$ is

$$\hat{p} - \hat{q} = \frac{1}{n} \sum_{i=1}^n X_i - \frac{1}{m} \sum_{i=1}^m Y_i$$

with estimated standard error

$$\hat{\text{se}}_{p-q} = \sqrt{\mathbb{V}(\hat{p} - \hat{q})} = \sqrt{\mathbb{V}(\hat{p}) + \mathbb{V}(\hat{q})} = \sqrt{\frac{\hat{p}(1 - \hat{p})}{n} + \frac{\hat{q}(1 - \hat{q})}{m}}$$

Since the Y_i 's are iid, by Central Limit Theorem \hat{q} is asymptotically normal with mean q and variance $\hat{\text{se}}_q^2$. The difference of two asymptotically normal random variables is asymptotically normal, thus $p - q$ is asymptotically normal with mean $p - q$ and variance $\hat{\text{se}}_{p-q}^2$. An approximate 90% confidence interval is

$$(\hat{p} - \hat{q} - 1.645\hat{\text{se}}_{p-q}, \hat{p} - \hat{q} + 1.645\hat{\text{se}}_{p-q})$$

7.9 An estimate for $p_1 - p_2$ is $0.9 - 0.85 = 0.05$ with standard error

$$\sqrt{\frac{0.9(1 - 0.9)}{100} + \frac{0.85(1 - 0.85)}{100}} = 0.0466368953$$

with 80% and 90% confidence intervals given by

$$\begin{aligned} 80\% : \quad & (\hat{p} - \hat{q} - 1.282\text{se}_{p-q}, \hat{p} - \hat{q} + 1.282\text{se}_{p-q}) = (-0.0097885, 0.1097885) \\ 90\% : \quad & (\hat{p} - \hat{q} - 1.96\text{se}_{p-q}, \hat{p} - \hat{q} + 1.96\text{se}_{p-q}) = (-0.041408315, 0.141408315) \end{aligned}$$

8.7 (a)

$$\begin{aligned} \mathbb{P}(\hat{\theta} \leq k) &= \mathbb{P}(\max\{X_1, \dots, X_n\} \leq k) \\ &= \prod_{i=1}^n \mathbb{P}(X_i \leq k) \\ &= \left(\frac{k}{\theta}\right)^n \end{aligned}$$

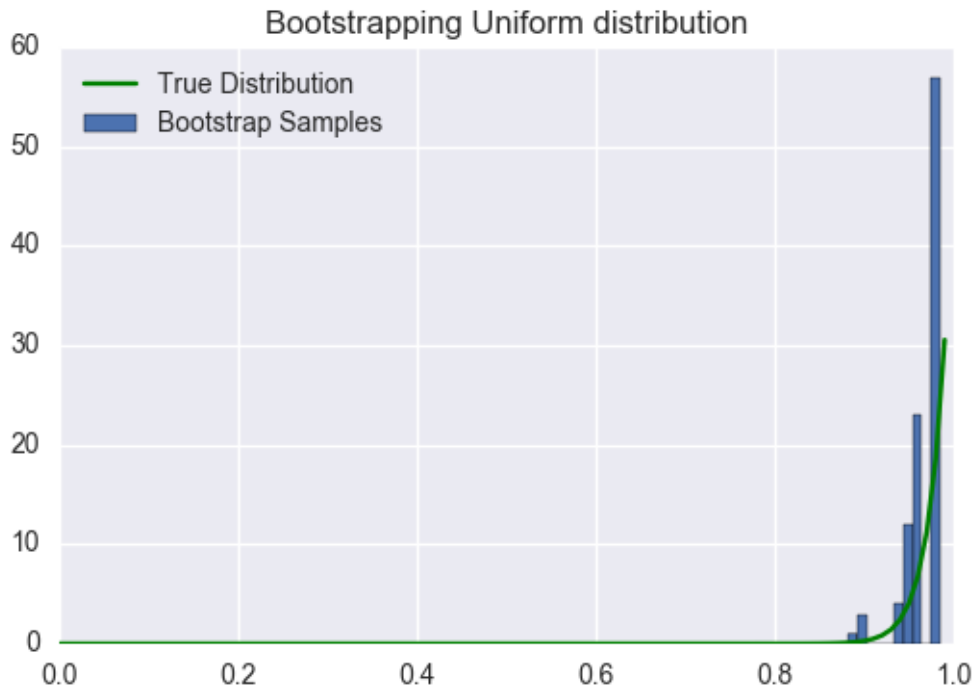


Figure 1: Comparison of the true distribution $\hat{\theta}$ to histograms from bootstrap

Code for the plot:

```
import numpy as np
import matplotlib.pyplot as plt
def sample(theta, n):
    """
    Draws n samples from uniform distribution in the interval (0, theta).
    """
    return np.random.uniform(0, theta, n)

def bootstrap(sample, B):
    """
    Performs bootstrapping B times from the given sample.
    """
    n = sample.shape[0]
    bootstrap = np.zeros((B, 50))
    for i in range(B):
        bootstrap[i, :] = np.random.choice(sample, 50)
```

```

    return bootstrap
def maxestimator(bootstrap):
    """
    Returns the maximum value from each bootstrap sample.
    """
    return np.max(bootstrap, axis = 1)
def plotcdf(theta, n):
    """
    Plots the true distribution of the X_max
    """
    x = np.arange(0, theta, 0.01)
    f = lambda x: n*(x/theta) ** (n-1)
    y = f(x)
    plt.plot(x, y, 'g', label = 'True Distribution')
def finalplot(theta, n, B):
    """
    Plots both the simulations and the true distribution for comparison.
    """
    samples = sample(theta, n)
    bootstraps = bootstrap(samples, B)
    max_samples = maxestimator(bootstraps)
    plt.hist(max_samples, label = 'Bootstrap Samples')
    plotcdf(1, 50)
    plt.legend(loc='best')
    plt.title('Bootstrapping Uniform distribution')
finalplot(1, 50, 100)

```

(b) Let $\hat{\theta} = X_{max} = \max\{X_1, \dots, X_n\}$. Then

$$\begin{aligned}\mathbb{P}(\hat{\theta}^* = \hat{\theta}) &= 1 - \mathbb{P}(\hat{\theta}^* \neq \hat{\theta}) \\ &= 1 - \left(1 - \frac{1}{n}\right)^n\end{aligned}$$

The second equality holds as $\mathbb{P}(\hat{\theta}^* \neq \hat{\theta})$ denotes the probability that any random sampling with replacement of the n samples drawn has probability of $1 - 1/n$ not being x_{max} (which is fixed since a random sample of n has been drawn). As each sampling process is iid due to replacement, probability of them all not being x_{max} is $(1 - 1/n)^n$. Thus we have $\mathbb{P}(\hat{\theta}^* = \hat{\theta}) \rightarrow 0$ as $n \rightarrow \infty$ and $\mathbb{P}(\hat{\theta}^* = \hat{\theta}) = .632$ for $n = 50$.

9.2 (a) For $X_1, \dots, X_n \sim \text{Uniform}(a, b)$

$$\begin{aligned}\hat{\mu} = \bar{X}_n &= \frac{1}{n} \sum_{i=1}^n X_i = \frac{\hat{b} + \hat{a}}{2} \\ \hat{\sigma}^2 &= \frac{1}{n} \sum_{i=1}^n (\bar{X}_n - X_i)^2 = \frac{(\hat{b} - \hat{a})^2}{12}\end{aligned}$$

thus

$$\begin{aligned}(\hat{b} - \hat{a}) + (\hat{b} + \hat{a}) &= +\sqrt{12\hat{\sigma}^2} + 2\hat{\mu} \\ \hat{b} &= \frac{1}{2} \left(\sqrt{12\hat{\sigma}^2} + 2\hat{\mu} \right) \\ \hat{a} &= 2\hat{\mu} - \hat{b}\end{aligned}$$

the positive root is taken as $b - a > 0$.

(b) Let $X_1, \dots, X_n \sim \text{Uniform}(a, b)$, with $X_{max} = \max\{X_1, \dots, X_n\}$. If $b < X_{max}$, then $f(X_j; a, b) = 0$ for some j . Thus if $b \geq X_{max}$, then $f(X_i; a, b) = 1/b - a$ for all i . In a similar fashion, letting $X_{min} = \{X_1, \dots, X_n\}$, if $X_{min} < a$ we also have $f(X_j; a, b) = 0$ for some j and $f(X_i; a, b) = 1/b - a$ for all i if $X_{min} \geq a$. Therefore,

$$\mathcal{L}_n(a, b) := \begin{cases} 0, & X_{min} < a \text{ or } X_{max} > b \\ \left(\frac{1}{b-a}\right)^n, & \text{otherwise} \end{cases}$$

$\mathcal{L}(a, b)$ strictly decreasing over $(-\infty, X_{\min}]$ and $[X_{\max}, \infty)$, thus the maximum likelihood estimators $\hat{a} = X_{\min}$ and $\hat{b} = X_{\max}$.

- (c) Let $\tau = \int x dF(x)$ be given, then from (b) we know that MLE's \hat{a} and \hat{b} are given by X_{\min} and X_{\max} respectively. Then the MLE of τ follows from MLE's \hat{a} and \hat{b} . Thus MLE of τ is $(X_{\min} + X_{\max})/2$.
- (d) By simulation, the MSE of $\hat{\tau} \approx 0.015$ by using the Python code below:

```
import numpy as np
n = 500000
mle_tau = np.zeros((n,1))
for i in np.arange(n):
    s = np.random.uniform(1, 3, 10)
    s_max = np.max(s)
    s_min = np.min(s)
    mle_tau[i] = 0.5 * (s_max + s_min)
(1/n) * np.sum((mle_tau - 2) **2)
```

Analytically, for the MSE of the nonparametric plugin estimator $\tilde{\tau}$ we have

$$\begin{aligned}\mathbb{E}(\hat{\theta} - \theta)^2 &= \mathbb{E}(\hat{\theta}^2 - 2\hat{\theta}\theta + \theta^2) \\ &= \mathbb{E}(\hat{\theta}^2) - 2\theta\mathbb{E}(\hat{\theta}) + \mathbb{E}(\theta^2) \\ &= \mathbb{E}(\hat{\theta}^2) - 2\theta\mathbb{E}(\hat{\theta}) + \mathbb{E}(\theta^2) \\ &= \mathbb{E}(\hat{\theta}^2) - 4\end{aligned}$$

$$\begin{aligned}\mathbb{E}(\hat{\theta})^2 &= n^{-2} \left[\mathbb{E} \left(\sum_{i=1}^n X_i^2 \right) + 2\mathbb{E} \left(\sum_{i \neq j} X_i X_j \right) \right] \\ &= n^{-2} [n\mathbb{E}(X^2) + n(n-1)\mathbb{E}(X_i X_j)] \\ &= n^{-2} [n\mathbb{E}(X^2) + n(n-1)\mathbb{E}(X)^2] \\ &= 121/30\end{aligned}$$

using the substitution $\mathbb{E}(X^2) = 2$, $\mathbb{E}(X) = 13/2$ and $n = 10$. The expectations are computed with $a = 1, b = 2$. Thus we have MSE to be $1/30$.

9.6 (a) The log-likelihood function is given to be

$$\begin{aligned}\ell_n(\theta) &= n \log \frac{1}{\sqrt{2\pi}} - \sum_{i=1}^n \frac{(X_i - \theta)^2}{2} \\ &= n \log \frac{1}{\sqrt{2\pi}} - \frac{n}{2} S^2 - \frac{n}{2} (\bar{X}_n - \theta)^2\end{aligned}$$

with $\frac{\partial \ell_n}{\partial \theta} = n(\bar{X}_n - \theta)$

where $\bar{X}_n = n^{-1} \sum_{i=1}^n X_i$. Thus the maximum of the log-likelihood function is when $\theta = \bar{X}_n$. Now, $X_1 \sim N(\bar{X}_n, 1)$ and so the maximum likelihood of ψ is given by:

$$\begin{aligned}\hat{\psi} &= \mathbb{P}(X_1 > 0) = \mathbb{P}(X_1 - \bar{X}_n > -\bar{X}_n) \\ &= \mathbb{P}(Z > -\bar{X}_n)\end{aligned}$$

where the Z refers to the standard normal distribution.

- (b) The distribution of $\hat{\psi}$ is normally distributed with $\hat{s}_e = \sqrt{1/I_n(\hat{\theta})}$

$$\begin{aligned}I_n(\theta) &= nI(\theta) \\ &= nI(\theta)\end{aligned}$$

Since $f(X; \theta) = 1/\sqrt{2\pi} \exp(-1/2(X - \theta)^2)$, we have the score function $s(X; \theta) = X - \theta$ and $s'(X; \theta) = 1$. Thus $I(\theta) = -\mathbb{E}_\theta(s'(X; \theta)) = -1$. Hence $\hat{\text{se}} = 1/\sqrt{n}$ which gives us a 95 percent confidence interval of $\mathbb{P}(Z > -\bar{X}_n) \pm 1.96/\sqrt{n}$

- (c) Let $\tilde{\psi} = (1/n) \sum_i Y_i$, to show $\tilde{\psi}$ is consistent, we need to show that $\mathbb{P}(|\tilde{\psi} - \psi| > \epsilon) \rightarrow 0$ as $n \rightarrow \infty$. We also observe that $Y \sim \text{Bernoulli}(p)$ where $p = \mathbb{P}(X > 0)$

$$\begin{aligned} \mathbb{P}(|\tilde{\psi} - \psi| > \epsilon) &= \mathbb{P}(|(1/n) \sum_i Y_i - \mathbb{P}(Y_1 = 1)| > \epsilon) \\ &= \mathbb{P}\left(\left|\sum_i Y_i - n\mathbb{P}(Y_1 = 1)\right| > (\epsilon n)^2\right) \\ &\leq \mathbb{E}\left(\left[\sum_i Y_i - n\mathbb{P}(Y_1 = 1)\right]^2\right) / (\epsilon n)^2 \quad \text{by Markov's inequality} \\ &= \frac{n\mathbb{E}(Y) - n\mathbb{E}(Y)^2 + (n\mathbb{E}(Y) - n\mathbb{P}(Y_1 = 1))^2}{\epsilon^2 n^2} \\ &\rightarrow \frac{(\mathbb{E}(Y) - \mathbb{P}(Y_1 = 1))^2}{\epsilon^2} = 0 \text{ as } n \rightarrow \infty \end{aligned}$$

- (d) We are given $\tilde{\psi} = \bar{Y}_n = n^{-1} \sum_{i=1}^n Y_i$ with $Y \sim \text{Bernoulli}(p)$, $p = \mathbb{P}(Z > -\theta)$. Thus by Central Limit Theorem,

$$\frac{\bar{Y}_n - \mathbb{E}(Y)}{\sqrt{\text{Var}(Y)/n}} \rightsquigarrow N(0, 1)$$

thus \bar{Y}_n is normally distributed with mean p and variance $p(1-p)/n$. Next we consider the MLE, $\hat{\psi}$. We know from (a) that $\psi(\theta) = \mathbb{P}(Z > -\theta)$ with $X \sim N(\theta, 1)$ and $\hat{\psi} = \psi(\bar{X}_n)$. By Central Limit Theorem,

$$\frac{\bar{X}_n - \theta}{1/\sqrt{n}} \rightsquigarrow N(0, 1)$$

and $\hat{\psi}$ is differentiable since it is a complementary cumulative distribution function and its derivative is strictly greater than 0. Therefore,

$$\frac{\psi(\bar{X}_n) - \psi(\theta)}{|\psi'(\bar{X}_n)|} \rightsquigarrow N(0, 1)$$

evaluating $\psi'(\bar{X}_n)$, we have the probability distribution function evaluated at \bar{X}_n which is

$$f_X(\bar{X}_n) = \frac{1}{\sqrt{2\pi}} \exp(-1/2(\bar{X}_n)^2)$$

Letting $f_Z(z)$ and $\bar{F}_Z(z)$ denote the pdf and complementary cdf of a normal distribution with mean 0 and variance 1, the asymptotic relative efficiency of $\tilde{\psi}$ to $\hat{\psi}$ is

$$\text{ARE}(\tilde{\psi}, \hat{\psi}) = \frac{\bar{F}_Z(-\theta)(1 - \bar{F}_Z(-\theta))}{(f_Z(\bar{X}_n))^2}$$

- (e) If the data is not normal, unless it is a constant distribution, it will be consistent due to Central Limit Theorem. Let $X_i = -1$ for all i , i.e. no matter how many times you sample you only get -1. Thus

$$\hat{\psi} = \mathbb{P}(Z > -1) \neq 0 = \mathbb{P}(Y_1 = 1) = \psi$$