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Statistics: Homework 2

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

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

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

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

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

For $X_1, \dots, X_n \sim \text{Bernoulli}(p)$ and $Y_1, \dots, Y_n \sim \text{Bernoulli}(q)$ 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

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

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

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

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

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

8.7 (a)

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

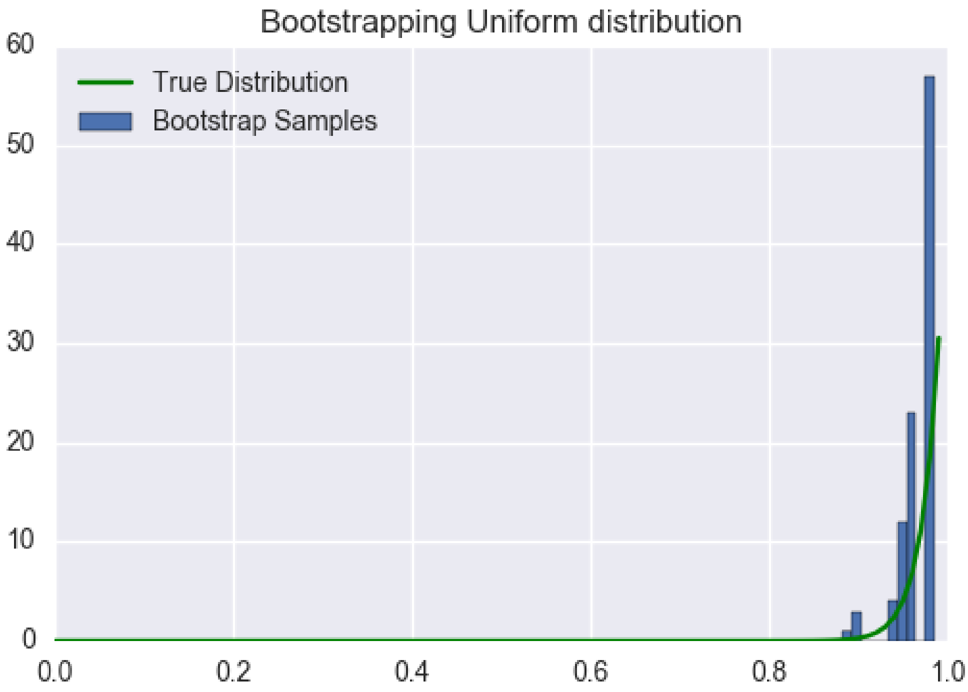

Figure 1: Comparison of the true distribution $\hat{\theta}$ to histrograms from bootstrap

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 $\mathbb{P}(\hat{\theta}^* = \hat{\theta}) \to 0$ as $n \to \infty$ and $\mathbb{P}(\hat{\theta}^* = \hat{\theta}) = .632$ for $n = 50$.

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

$$\hat{\mu} = \overline{X}_n = \frac{1}{n}\sum_{i=1}^{n} X_i = \frac{\hat{b} + \hat{a}}{2}$$

$$\hat{\sigma}^2 = \frac{1}{n}\sum_{i=1}^{n} (\overline{X}_n - X_i)^2 = \frac{(\hat{b} - \hat{a})^2}{12}$$

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:

```python
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

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

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

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

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

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

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

where the Z refers to the standard normal distribution.

(b) By Central Limit Theorem, we know that

$$\sqrt{n}(\overline{X}_n - \theta) \rightsquigarrow N(\theta, 1)$$

defining $g(\theta) = \psi = \mathbb{P}(Y_1 = 1) = \mathbb{P}(X_1 > 0)$ we have

$$\frac{g(\overline{X}_n) - g(\theta)}{|g'(\overline{X}_n)|\widehat{\mathsf{se}}(\overline{X}_n)} \rightsquigarrow N(0, 1)$$

thus the $\hat{\mathsf{se}}$ for the confidence interval is $|g'(\overline{X}_n)|\hat{\mathsf{se}} = \frac{1}{\sqrt{2n\pi}} \exp(-1/2(\overline{X}_n)^2)$. Thus the confidence interval is given by:

$$\mathbb{P}\left(Z > -\overline{X}_n\right) \pm 1.96 \frac{1}{\sqrt{2n\pi}} \exp(-1/2(\overline{X}_n)^2)$$

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

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

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

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

thus $\overline{Y}_n$ 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 $\hat{\psi} = \psi(\overline{X}_n)$. By Central Limit Theorem,

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

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

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

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

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

Letting $f_Z(z)$ and $\overline{F}_Z(z)$ 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

$$\mathsf{ARE}(\tilde{\psi}, \hat{\psi}) = \frac{\overline{F}_Z(-\theta)(1 - \overline{F}_Z(-\theta))}{\left(f_Z(\overline{X}_n)\right)^2}$$

(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

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