STA261 - Module 5

Asymptotic Extensions

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Limitations of Finite Sample Sizes

- In almost everything we've done so far, we've assumed a sample $X_1, X_2, \ldots, X_n \overset{iid}{\sim} f_\theta$ of fixed size n
- We've needed to know the distributions of various statistics of $X_1, X_2, \ldots, X_n$
- This requirement has been very limiting, as the distributions of most statistics don't have closed forms (or are unknown entirely)

eg, $\sum_{i=1}^{n} a_i \cdot \chi^2_{(i)}$ has no closed form, $a_i \in \mathbb{R}$.

- Even the exact distribution of the sample mean $\frac{1}{n}\sum_{i=1}^{n} X_i$ is only available for a few parametric families

even though we use $\bar{X}_n$, like, EVERYWHERE!

On the other hand, $\bar{X}_n \xrightarrow{d} \mathbb{E}(X_i)$ (assuming iid, $\mathbb{E}[X] < \infty$, etc)

Driving Up the Sample Size

- On the other hand, we have plenty of *limiting* distributions as $n \to \infty$
- Example 5.1: $X_1, \ldots X_n \overset{iid}{\sim} N(\mu, \sigma^2) \Rightarrow \bar{X}_n \xrightarrow{p} \mu$ by LLN, $\frac{\bar{X}_n - \mu}{\sqrt{\sigma^2/n}} \longrightarrow N(0,1)$
- Example 5.2: $X_n \sim Bin(n, p_n)$ with $n \cdot p_n \overset{n\to\infty}{\to} \lambda$, then $X_n \xrightarrow{d} Poisson(\lambda)$
- Of course, we never have $n = \infty$ in real life
- But if we have the luxury of a very large sample size, the "difference" between the exact distribution and the limiting distribution should (hopefully) be tolerable
- Since the Normal distribution is particularly nice, we will milk the CLT for all it's worth

A Review of Standard Limiting Results

$X_n \xrightarrow{d} X$ means that $F_{X_n}(x) \to F_X(x)$ at all continuity points of $F_X(\cdot)$.
$X_n \xrightarrow{p} X$ means that $\forall \varepsilon > 0$, $P(|X_n - X| \geq \varepsilon) \to 0$
$X_n \xrightarrow{a.s.} X$ means that $P(\lim_{n\to\infty} X_n = X) = 1$ (not used in our course)

- In the following, let $\{X_n\}_{n\geq 1}$ and $\{Y_n\}_{n\geq 1}$ be sequences of random variables, let X be another random variable, let $x, y \in \mathbb{R}$ be constants, and let $g(\cdot)$ be a continuous function

- Theorem 5.1: If $X_n \xrightarrow{p} X$, then $X_n \xrightarrow{d} X$. If $X_n \xrightarrow{d} x$, then $X_n \xrightarrow{p} x$.

Converse is not true in general, except when

- Theorem 5.2 (**Slutsky's theorem**): If $X_n \xrightarrow{d} X$ and $Y_n \xrightarrow{p} y$, then $Y_n \cdot X_n \xrightarrow{d} y \cdot X$ and $X_n + Y_n \xrightarrow{d} X_n + y$.

- Theorem 5.3 (**Continuous mapping theorem**): If $X_n \xrightarrow{p} X$, then $g(X_n) \xrightarrow{p} g(X)$. If $X_n \xrightarrow{d} X$, then $g(X_n) \xrightarrow{d} g(X)$. Also a.s. convergence!

Poll Time!

$\bar{X}_n$ sample mean of $N(3,1)$

$\bar{Y}_n$ sample mean of $N(5,1)$

a) $\bar{X}_n^2 - \bar{Y}_n^2 \longrightarrow -16$ in d but not p

b) $\bar{X}_n^2 - \bar{Y}_n^2 \longrightarrow -16$ in p but not d

c) $\bar{X}_n^2 - \bar{Y}_n^2 \longrightarrow -16$ in d and p

d) None of the above

$\bar{X}_n \xrightarrow{p} \mathbb{E}[X_i] = 3$ by WLLN

$\Rightarrow \bar{X}_n^2 \xrightarrow{p} 9$ by CMT

$\bar{Y}_n \xrightarrow{p} \mathbb{E}[Y_i] = 5$

$\Rightarrow \bar{Y}_n^2 \xrightarrow{p} 25$

$\Rightarrow \bar{X}_n^2 - \bar{Y}_n^2 \xrightarrow{d} 9 - 25 = -16$ by Slutsky

$\Rightarrow \bar{X}_n^2 - \bar{Y}_n^2 \xrightarrow{p} -16$ by Thm 5.1

Notation Update

- For the rest of this module, we will accentuate statistics of finite samples with the subscript n (so $\mathbf{X}$ is now $\mathbf{X}_n$, $\bar{X}$ is now $\bar{X}_n$, and so on)
- For a generic statistic, we'll write $T_n = T_n(\mathbf{X}_n)$
- If we're talking about a limiting property of a sequence $\{T_n\}_{n\geq 1}$, we'll abuse notation and just write that T_n has that limiting property, when the meaning is clear from context
- Example 5.3: Instead of "the sequence of sample means $\{\overline{X}_n\}_{n\geq 1}$ converges in probability to $\mathbb{E}[X_i]$"

  we'll say "$\overline{X}_n$ converges in probability to $\mathbb{E}[X_i]$"

  or "$\overline{X}_n \xrightarrow{p} \mathbb{E}[X_i]$"

Two Big Ones

- Theorem 5.4 (**Weak law of large numbers (WLLN)**): Let $X_1, X_2, \ldots$ be a sequence of iid random variables with $\mathbb{E}[X_i] = \mu$. Then

$$\bar{X}_n \xrightarrow{p} \mu.$$

(Also the Strong Law: $\bar{X}_n \xrightarrow{a.s.} \mu$)

- Theorem 5.5 (**Central limit theorem (CLT)**): Let $X_1, X_2, \ldots$ be a sequence of iid random variables with $\mathbb{E}[X_i] = \mu$ and $\text{Var}(X_i) = \sigma^2$. Then

$$\frac{\bar{X}_n - \mu}{\sqrt{\sigma^2/n}} \xrightarrow{d} \mathcal{N}(0,1).$$

- The CLT is equivalent to $\sqrt{n}(\bar{X}_n - \mu) \xrightarrow{d} \mathcal{N}(0, \sigma^2)$, which is the form we'll be using most often

(by Slutsky!)

Asymptotic Unbiasedness

- As in Module 2, we're interested in estimators of $\tau(\theta)$
- But now we're concerned with their limiting behavious as $n \to \infty$
- For finite n, we insisted that our "best" estimators be unbiased
- In the asymptotic setup, we can relax that slightly
- Definition 5.1: Suppose that $\{W_n\}_{n\geq 1}$ is a sequence of estimators for $\tau(\theta)$. If $\text{Bias}_\theta(W_n) \xrightarrow{n\to\infty} 0$ for all $\theta \in \Theta$, then $\{W_n\}_{n\geq 1}$ is said to be **asymptotically unbiased** for $\tau(\theta)$.
- Example 5.4: In the $N(\mu, \sigma^2)$ model, $\frac{1}{n+1}\sum_{i=1}^n X_i$ is asymptotically unbiased for μ.

  $\mathbb{E}_\mu\left[\frac{1}{n+1}\sum_{i=1}^n X_i\right] = \frac{n}{n+1}\mu$, so $\text{Bias}_\mu\left(\frac{1}{n+1}\sum_{i=1}^n X_i\right) = \mu\cdot\left(\frac{n}{n+1}-1\right) \xrightarrow{n\to\infty} 0$

Consistency

- $\overline{X}_n \xrightarrow{p} \mu$ is the prototypical example of an estimator converging in probability to the "right thing"
- We have a special name for this
- Definition 5.2: A sequence of estimators W_n of $\tau(\theta)$ is said to be **consistent** if $W_n \xrightarrow{p} \tau(\theta)$ for every $\theta \in \Theta$. for $\tau(\theta)$
- Example 5.5:

$\text{Exp}(\lambda)$: $1/\overline{X}_n^2$ is consistent for λ^2. Why? $\overline{X}_n \xrightarrow{p} 1/\lambda$ by WLLN.

$\Rightarrow g(\overline{X}_n) \xrightarrow{p} g(1/\lambda)$ by CMT, where $g(x) = 1/x^2$

$\Rightarrow 1/\overline{X}_n \xrightarrow{p} \lambda^2$. So yep.

$N(\mu, \sigma^2)$: $\dfrac{\overline{X}_n^2}{\overline{X}_n^2}$ is consistent for $\dfrac{\mu^2}{\mu^2+\sigma^2}$.

Why? $\overline{X}_n \xrightarrow{p} \mu$ by WLLN

$\Rightarrow \overline{X}_n^2 \xrightarrow{p} \mu^2$ by CMT

$\overline{X}_n^2 = \frac{1}{n}\sum_{i=1}^n X_i^2 \xrightarrow{p} \mathbb{E}[X_i^2] = \mu^2 + \sigma^2$ by WLLN

By Slutsky, $\dfrac{\overline{X}_n^2}{\overline{X}_n^2} \xrightarrow{d} \dfrac{\mu^2}{\mu^2+\sigma^2}$. By Theorem 5.1, $\dfrac{\overline{X}_n^2}{\overline{X}_n^2} \xrightarrow{p} \dfrac{\mu^2}{\mu^2+\sigma^2}$.

Showing Consistency

- Sometimes it's easy to show consistency directly from the definition

- Example 5.6: Let $X_1, X_2, \ldots, X_n \overset{iid}{\sim} \mathcal{N}\left(\mu, \sigma^2\right)$, where $\mu \in \mathbb{R}$ and $\sigma^2 > 0$. Is the sample mean $\overline{X}_n$ consistent for μ?

Let $\varepsilon > 0$. Then $\mathbb{P}_\mu(|\overline{X}_n - \mu| < \varepsilon)$

$= \mathbb{P}_\mu(-\varepsilon < \overline{X}_n - \mu < \varepsilon)$

$= \mathbb{P}_\mu\left(\frac{-\varepsilon}{\sqrt{\sigma^2/n}} < \frac{\overline{X}_n - \mu}{\sqrt{\sigma^2/n}} < \frac{\varepsilon}{\sqrt{\sigma^2/n}}\right)$

$= \mathbb{P}_\mu\left(\frac{-\varepsilon}{\sqrt{\sigma^2/n}} < Z < \frac{\varepsilon}{\sqrt{\sigma^2/n}}\right)$ where $Z \sim N(0,1)$

$= \Phi\left(\frac{\varepsilon\sqrt{n}}{\sigma}\right) - \Phi\left(\frac{-\varepsilon\sqrt{n}}{\sigma}\right)$

$\xrightarrow{n \to \infty} \Phi(\infty) - \Phi(-\infty) = 1.$

$\Rightarrow \overline{X}_n \xrightarrow{P} \mu.$

Showing Consistency

- It's usually easier to use standard limiting results (Slutsky, continuous mapping, etc.) than to go directly from the definition

- Example 5.7: Let $X_1, X_2, \ldots, X_n \overset{iid}{\sim} \mathcal{N}(\mu, \sigma^2)$, where $\mu \in \mathbb{R}$ and $\sigma^2 > 0$. Is the sample variance S_n^2 consistent for σ^2? Exercise: try doing this from the definition!

$$S_n^2 = \frac{1}{n-1}\sum_{i=1}^{n}(X_i - \bar{X}_n)^2$$

$$= \frac{n}{n-1}\left[\frac{1}{n}\sum_{i=1}^{n}(X_i - \bar{X}_n)^2\right]$$

$$= \frac{n}{n-1}\left[\frac{1}{n}\sum_{i=1}^{n}(X_i - \mu)^2 + (\bar{X}_n - \mu)^2\right] \xrightarrow{d} 1 \cdot (\sigma^2 + 0) = \sigma^2 \quad \text{by Slutsky (x2)}$$

$$\Rightarrow S_n^2 \xrightarrow{p} \sigma^2 \text{ by Theorem 5.1.}$$

① $\xrightarrow{p} 1$

② $= \overline{(X-\mu)^2} \xrightarrow{p} \mathbb{E}[(X-\mu)^2] = \sigma^2$ by LLN

③ $= (\overline{X-\mu})^2 \xrightarrow{p} \mathbb{E}[X-\mu]^2 = 0$ by LLN + CMT

Bringing Back the MSE

- In Module 2, we compared estimators by their MSEs
- To extend that idea to the asymptotic setup, we need a new mode of convergence
- Definition 5.3: Suppose that W_n is a sequence of estimators for $\tau(\theta)$. If $\text{MSE}_\theta(W_n) \xrightarrow{n\to\infty} 0$ for all $\theta \in \Theta$, then W_n is said to **converge in MSE** ~~for~~ to $\tau(\theta)$. "$W_n \xrightarrow{\text{MSE}} \tau(\theta)$."
- Example 5.8: $X_1, \dots, X_n \overset{iid}{\sim} \text{Bin}(k, p)$, $p \in (0,1)$. Then $\bar{X}_n \xrightarrow{\text{MSE}} p$, because

$$\begin{aligned}
\text{MSE}_p(\bar{X}_n) &= \underbrace{\text{Bias}_p(\bar{X}_n)^2}_{0} + \text{Var}_p(\bar{X}_n) \\
&= \text{Var}_p(\bar{X}_n) \\
&= \frac{1}{n} kp(1-p) \xrightarrow{n\to\infty} 0 \quad \forall p \in (0,1).
\end{aligned}$$

So $\bar{X}_n \xrightarrow{\text{MSE}} p$.

Poll Time!

$$MSE_\theta[g(W_n)] = \mathbb{E}_\theta[(g(W_n) - g(\theta))^2]$$
$$= Bias_\theta(g(W_n))^2 + Var_\theta(g(W_n))$$

$$MSE_\theta(W_n) = Bias_\theta(W_n)^2 + Var_\theta(W_n)$$
$$= Var_\theta(W_n)$$
$$\downarrow n\to\infty \iff \downarrow n\to\infty$$
$$0 \qquad 0.$$

Convergence in MSE is Already Good Enough

- It turns out that convergence in MSE is strong enough to guarantee consistency

- Theorem 5.6: If W_n is a sequence of estimators for $\tau(\theta)$ that converges in MSE for all $\theta \in \Theta$, then W_n is consistent for $\tau(\theta)$.

Proof. Let $\varepsilon > 0$, and let $\theta \in \Theta$. By Chebyshev's inequality,

$$P_\theta(|W_n - \tau(\theta)| \geq \varepsilon) \leq \frac{E_\theta[(W_n - \tau(\theta))^2]}{\varepsilon^2} = \frac{MSE_\theta(W_n)}{\varepsilon^2} \xrightarrow{n \to \infty} \frac{0}{\varepsilon^2} = 0.$$

So $W_n \xrightarrow{P} \tau(\theta)$, and so W_n is consistent for $\tau(\theta)$. □

A Criterion for Consistency

- If we know $\mathbb{E}_\theta[W_n]$ and $\text{Var}_\theta(W_n)$, this next theorem often makes short work out of checking for consistency

- Theorem 5.7: If W_n is a sequence of estimators for $\tau(\theta)$ such that $\text{Bias}_\theta(W_n) \xrightarrow{n\to\infty} 0$ and $\text{Var}_\theta(W_n) \xrightarrow{n\to\infty} 0$ for all $\theta \in \Theta$, then W_n is consistent for $\tau(\theta)$.

Proof. For any $\theta \in \Theta$, $MSE_\theta(W_n) = \text{Bias}_\theta(W_n)^2 + \text{Var}_\theta(W_n)$

By assumption, $\text{Bias}_\theta(W_n) \xrightarrow{n\to\infty} 0$ and $\text{Var}_\theta(W_n) \xrightarrow{n\to\infty} 0$.

So $MSE_\theta(W_n) \xrightarrow{n\to\infty} 0$.

By Theorem 5.6, W_n is consistent for θ. □

The Sample Mean is Always Consistent

- Example 5.9: Let $X_1, X_2, \ldots, X_n \overset{iid}{\sim} f_\theta$, where $\mathbb{E}[X_i] = \mu$. Show that $\overline{X}_n$ is consistent for μ.

By WLLN, $\overline{X}_n \xrightarrow{p} \mathbb{E}[X_i] = \mu$.

So $\overline{X}_n$ is consistent for μ.

The Sample Variance is Always Consistent

- One can (very tediously) show that if $X_1, X_2, \ldots, X_n$ are a random sample from a distribution with a finite fourth moment, then

$$\text{Var}\left(S_n^2\right) = \frac{\mathbb{E}\left[(X_i - \mathbb{E}\left[X_i\right])^4\right]}{n} - \frac{\text{Var}\left(X_i\right)^2 (n-3)}{n(n-1)}$$

- Example 5.10: Let $X_1, X_2, \ldots, X_n \overset{iid}{\sim} f_\theta$, where $\mathbb{E}\left[X_i\right] = \mu$ and $\text{Var}\left(X_i\right) = \sigma^2$. Show that S_n^2 is consistent for σ^2.

and f_θ has a finite fourth moment.

$\text{Bias}_{\sigma^2}(S_n^2) = 0$ from Assignment 0

$\text{Var}_{\sigma^2}(S_n^2) = \underbrace{\frac{\mathbb{E}[(X_i-\mu)^4]}{n}}_{\to 0 \text{ as } n\to\infty} - \underbrace{\frac{\sigma^4(n-3)}{n(n-1)}}_{\to 0 \text{ as } n\to\infty} \xrightarrow{n\to\infty} 0$

By Theorem 5.7, S_n^2 is consistent for σ^2.

Choosing Among Consistent Estimators

- Consistency is practically the bare minimum we can ask for from a sequence of estimators
- There are usually plenty of sequences that are consistent for $\tau(\theta)$

Assignment 5: tons of examples to practice with

- Which one should we use?
- It's tempting to go with whichever has the lowest variance for fixed n, but that would rule out a lot of fine estimators
- Example 5.11: $X_1, \ldots, X_n$ iid Poisson(λ), $\lambda > 0$.

$\bar{X}_n$ and S_n^2 are both consistent for λ, by our previous examples.

We know that for fixed n, $\bar{X}_n$ is the UMVUE (from Module 2), but should we just ignore S_n^2?

– $X_1, \ldots, X_n \sim N(\mu, \sigma^2)$. S_n^2 and $\hat{\sigma}^2_{mle} = \frac{1}{n}\sum_{i=1}^{n}(X_i - \bar{X})^2$ are both consistent for σ^2. Which to choose?

Asymptotic Normality

- There's a much more useful criterion, but first we need an important CLT-inspired definition

- Definition 5.4: Let T_n be a sequence of estimators for $\tau(\theta)$. If there exists some $\sigma^2 > 0$ such that

$$\sqrt{n}[T_n - \tau(\theta)] \xrightarrow{d} \mathcal{N}\left(0, \sigma^2\right),$$

  then T_n is said to be **asymptotically normal** with mean $\tau(\theta)$ and **asymptotic variance** σ^2.

- By virtue of the CLT, most unbiased estimators are asymptotically normal

Why not just talk about the distribution of T_n itself as $n \to \infty$?

Usually, it's a constant: $\bar{X}_n \xrightarrow{d} \mu$, for example.

Asymptotic Normality: Examples

- Example 5.12: Let $X_1, X_2, \ldots, X_n \overset{iid}{\sim} \text{Bin}(k, p)$. Show that the sample mean $\overline{X}_n$ is asymptotically normal.

By the CLT, $\sqrt{n}(\overline{X}_n - \mathbb{E}_p(X_i)) \xrightarrow{d} N(0, \text{Var}_p(X_i))$

$\Rightarrow \sqrt{n}(\overline{X}_n - kp) \xrightarrow{d} N(0, kp(1-p))$

So $\overline{X}_n$ is asymptotically normal with mean kp and asymptotic variance $kp(1-p)$.

Asymptotic Normality: Examples

- Example 5.13: Let $X_1, X_2, \ldots, X_n \overset{iid}{\sim} \text{Exp}(\lambda)$. Show that the second sample moment ~~$\bar{X}_n$~~ $\overline{X^2}_n$ is asymptotically normal.

$$\mathbb{E}_\lambda[X_i^2] = \text{Var}_\lambda(X_i) + \mathbb{E}_\lambda[X_i]^2 = 2/\lambda^2$$

$$\text{Var}_\lambda(X_i^2) = \mathbb{E}_\lambda[X_i^4] - \mathbb{E}_\lambda[X_i^2]^2$$

$$= \frac{4!}{\lambda^4} - \left(\frac{2}{\lambda^2}\right)^2 = \frac{20}{\lambda^4}$$

use mgfs!

By the CLT, $\sqrt{n}\left(\overline{X^2}_n - 2/\lambda^2\right) \xrightarrow{d} N\left(0, \frac{20}{\lambda^4}\right)$.

So $\overline{X^2}_n$ is asymptotically normal w/ mean $2/\lambda^2$ and asymptotic variance $\frac{20}{\lambda^4}$.

Asymptotic Distributions

- More generally, we can talk about the limiting distribution of $\sqrt{n}[T_n - \tau(\theta)]$ even when it's not Normal

- Definition 5.5: Suppose that T_n is a sequence of estimators for $\tau(\theta)$. When it exists, the distribution of $\lim_{n\to\infty} \sqrt{n}[T_n - \tau(\theta)]$ is called the **asymptotic distribution** (or **limiting distribution**) of T_n.

- So if T_n is an asymptotically normal sequence of estimators for $\tau(\theta)$ with asymptotic variance σ^2, then its asymptotic distribution is $\mathcal{N}\left(0, \sigma^2\right)$

- Example 5.14: From Example 5.12, $\bar{X}_n$ has asymptotic distribution $N(0, kp(1-p))$.

- We might prefer to speak of the distribution of T_n itself when n is large

We can say "for large n, the dist. of $\bar{X}_n$ approaches $N(kp, \frac{kp(1-p)}{n})$"

$\sqrt{n}(\bar{X}_n - kp) \overset{\cdot}{\sim} N(0, kp(1-p))$

$\Rightarrow \bar{X}_n \overset{\cdot}{\sim} N(kp, \frac{kp(1-p)}{n})$

But we can't say "for large n, the dist. of $\bar{X}_n$ is $N(kp, \frac{kp(1-p)}{n})$"

Poll Time!

$$\bar{X}_n \xrightarrow{L} \mathbb{E}_\theta[X_i] = \theta.$$

The Delta Method

- If some sequence T_n is asymptotically normal for θ and some function $g(\cdot)$ is nice enough, then the next result gives a remarkably easy method of producing an asymptotically normal sequence of estimators of for $g(\theta)$

- Theorem 5.8 (**Delta method**): Suppose that $\theta \in \Theta \subseteq \mathbb{R}$ and $\sqrt{n}(T_n - \theta) \xrightarrow{d} \mathcal{N}\left(0, \sigma^2\right)$. If $g : \mathbb{R} \to \mathbb{R}$ is continuously differentiable with $g'(\theta) \neq 0$, then

$$\sqrt{n}[g(T_n) - g(\theta)] \xrightarrow{d} \mathcal{N}\left(0, [g'(\theta)]^2\sigma^2\right).$$

Assignment 5: way to handle $g'(\theta) = 0$

Proof. Taylor expand $g(T_n)$ around θ to get $g(T_n) = g(\theta) + g'(\tilde{\theta}_n)\cdot[T_n - \theta]$ for some $\tilde{\theta}_n$ between T_n and θ.

$$\Rightarrow \sqrt{n}(g(T_n) - g(\theta)) = \underbrace{g'(\tilde{\theta}_n)}_{①} \cdot \underbrace{\sqrt{n}[T_n - \theta]}_{②}$$

①: $T_n \xrightarrow{p} \theta$ by Slutsky, we must have $\tilde{\theta}_n \xrightarrow{p} \theta$ too

By CMT, $g'(\tilde{\theta}_n) \xrightarrow{p} g(\theta)$

②: $\sqrt{n}(T_n - \theta) \xrightarrow{d} N(0, \sigma^2)$ by assumption

By Slutsky,

$\sqrt{n}(g(T_n) - g(\theta)) \xrightarrow{d} g'(\theta)\cdot N(0, \sigma^2)$

$= N(0, [g'(\theta)]^2\cdot\sigma^2)$. □

The Delta Method: Examples

- Example 5.15: Let $X_1, X_2, \ldots, X_n \overset{iid}{\sim} \mathcal{N}\left(\mu, \sigma^2\right)$ where $\mu \in \mathbb{R} \setminus \{0\}$ and $\sigma^2 > 0$. Find the limiting distribution of $1/\overline{X}_n$.

Let $g(x) = 1/x$, so $g'(x) = -1/x^2$ for $x \neq 0$.

By CLT, $\sqrt{n}(\overline{X}_n - \mu) \xrightarrow{d} N(0, \sigma^2)$.

By the delta method, $\sqrt{n}(g(\overline{X}_n) - g(\mu)) \xrightarrow{d} N(0, [g'(\mu)]^2 \cdot \sigma^2)$

$\Rightarrow \sqrt{n}(1/\overline{X}_n - 1/\mu) \xrightarrow{d} N(0, \sigma^2/\mu^4)$.

So $1/\overline{X}_n$ has limiting distribution $N(0, \sigma^2/\mu^4)$.

So for large n, the distribution of $1/\overline{X}_n$ approaches $N(1/\mu, \frac{\sigma^2}{n\mu^4})$

The Delta Method: Examples

- Example 5.16: Let $X_1, X_2, \ldots, X_n \overset{iid}{\sim}$ Bernoulli (θ) where $\theta \in (0,1)$. ~~Find the~~ Find the limiting distribution of $\log\left(1 - \overline{X}_n\right)$.

Let $g(x) = \log(1-x)$, $x \in (0,1)$.

$\Rightarrow g'(x) = \frac{-1}{1-x}$, $x \in (0,1)$.

By the CLT, $\sqrt{n}(\overline{X}_n - \theta) \xrightarrow{d} N(0, \theta(1-\theta))$.

By the delta method, $\sqrt{n}(\log(1-\overline{X}_n) - \log(1-\theta)) \xrightarrow{d} N\left(0, \left(\frac{-1}{1-\theta}\right)^2 \theta(1-\theta)\right)$

$= N\left(0, \frac{\theta}{1-\theta}\right)$

So $\log(1-\overline{X}_n)$ has asymptotic distribution $N\left(0, \frac{\theta}{1-\theta}\right)$.

The Delta Method: Examples

- Example 5.17: Let $X_1, X_2, \ldots, X_n \overset{iid}{\sim} f_\theta$ where $\mathbb{E}_\theta[X_i] = \mu$ and $\text{Var}_\theta(X_i) = \sigma^2$. If $\tau : \mathbb{R} \to \mathbb{R}$ is continuously differentiable with $\tau'(\mu) \neq 0$, describe the distribution of $\tau(\overline{X}_n)$ as n becomes large.

By the CLT, $\sqrt{n}(\bar{X}_n - \mu) \xrightarrow{d} N(0, \sigma^2)$

By the delta method, $\sqrt{n}(\tau(\bar{X}_n) - \tau(\mu)) \xrightarrow{d} N(0, [\tau'(\mu)]^2 \cdot \sigma^2)$,

which is the asymptotic distribution of $\tau(\bar{X}_n)$.

So as n gets large, the distribution of $\tau(\bar{X}_n)$ approaches $N(\tau(\mu), [\tau'(\mu)]^2 \cdot \sigma^2/n)$.

Back to Choosing Estimators

- We know that when $T_n = \overline{X}_n$, the CLT says that

$$\frac{T_n - \mathbb{E}_\theta[T_n]}{\sqrt{\text{Var}_\theta(T_n)}} \xrightarrow{d} \mathcal{N}(0,1)$$

- Recall the Fisher information $I_n(\theta) = \text{Var}_\theta(S(\theta \mid \mathbf{X}_n))$ = $\mathbb{E}_\theta[S(\theta|\mathbf{X}_n)^2]$

- In Module 2, we said that an unbiased estimator W_n of $\tau(\theta)$ was efficient if its variance attained the Cramér-Rao Lower Bound $[\tau'(\theta)]^2/I_n(\theta)$

- We also noticed that if the X_i's were iid, then $I_n(\theta) = nI_1(\theta)$ by Theorem 2.10, under the conditions of the CRLB

Asymptotic Efficiency

- So if we could replace the T_n in the CLT statement with a general unbiased and efficient W_n, it would look like

redundant

$$\frac{W_n - \tau(\theta)}{\sqrt{[\tau'(\theta)]^2/nI_1(\theta)}} \xrightarrow{d} \mathcal{N}(0,1)$$

$= \mathbb{E}_\theta[W_n]$ since unbiased

$= \mathrm{Var}_\theta(W_n)$ since efficient

- Or equivalently

$$\sqrt{n}[W_n - \tau(\theta)] \xrightarrow{d} \mathcal{N}\left(0, \frac{[\tau'(\theta)]^2}{I_1(\theta)}\right)$$

- This is not a *result*, but a *condition* that we can demand of our estimators

- Definition 5.6: A sequence of estimators W_n is **asymptotically efficient** for $\tau(\theta)$ if

$$\sqrt{n}[W_n - \tau(\theta)] \xrightarrow{d} \mathcal{N}\left(0, \frac{[\tau'(\theta)]^2}{I_1(\theta)}\right)$$

If $\tau(\theta) = \theta$, then $\sqrt{n}(W_n - \theta) \xrightarrow{d} N(0, I_1^{-1}(\theta))$

Asymptotic Efficiency: Examples

- Example 5.18: Let $X_1, X_2, \ldots, X_n \overset{iid}{\sim} \text{Exp}(\lambda)$, where $\lambda > 0$. Show that $1/\overline{X}_n$ is asymptotically efficient for λ.

By the CLT, $\sqrt{n}(\overline{X}_n - 1/\lambda) \xrightarrow{d} N(0, 1/\lambda^2)$

By the delta method, $\sqrt{n}(1/\overline{X}_n - \lambda) \xrightarrow{d} N(0, \lambda^2)$.

$g(x) = 1/x, \quad x \neq 0$

$\Rightarrow g'(x) = -1/x^2, \quad x \neq 0$

$\Rightarrow g'(1/\lambda) = -\lambda^2$

$\Rightarrow [g'(1/\lambda)] = \lambda^4$

What's $I_1(\lambda)$? $l(\lambda|x) = \log(\lambda) - \lambda x$

$\Rightarrow S(\lambda|x) = \frac{1}{\lambda} - x$

$\Rightarrow -\frac{\partial}{\partial \lambda} S(\lambda|x) = 1/\lambda^2$

$\Rightarrow I_1(\lambda) = \mathbb{E}_x\left[-\frac{\partial}{\partial \lambda} S(\lambda|X)\right] = 1/\lambda^2$

Same thing!

So $1/\overline{X}_n$ is asymptotically efficient for λ.

Now $\frac{[\tau'(\lambda)]^2}{I_1(\lambda)} = \frac{1}{1/\lambda^2} = \lambda^2$

Asymptotic Efficiency: Examples

- Example 5.19: Let $X_1, X_2, \ldots, X_n \overset{iid}{\sim} \text{Poisson}(\lambda)$, where $\lambda > 0$. Show that $\overline{X}_n$ is asymptotically efficient for λ.

By the CLT, $\sqrt{n}(\overline{X}_n - \lambda) \xrightarrow{d} N(0, \lambda)$

$\tau(\lambda) = \lambda \Rightarrow [\tau'(\lambda)]^2 = 1.$

$L(\lambda|x) = \dfrac{e^{-\lambda}\lambda^x}{x!}$

$\Rightarrow \ell(\lambda|x) = -\lambda + x\cdot\log(\lambda) + c$, c free of λ

$\Rightarrow S(\lambda|x) = -1 + \dfrac{x}{\lambda}$

$\Rightarrow -\dfrac{\partial}{\partial\lambda} S(\lambda|x) = \dfrac{x}{\lambda^2}$

$\Rightarrow I_1(\lambda) = \mathbb{E}_\lambda\left[-\dfrac{\partial}{\partial\lambda} S(\lambda|X)\right] = \dfrac{1}{\lambda^2}\cdot\mathbb{E}_\lambda(X) = 1/\lambda$

So the asymptotic variance of $\overline{X}_n$ is $\lambda = \dfrac{[\tau'(\lambda)]^2}{I_1(\lambda)}$.

So $\overline{X}_n$ is asymptotically efficient for λ.

Large Sample Behaviour for the MLE

- We're ready to see why the MLE is almost always the point estimator of choice when n is large
- To understand this, we need to distinguish between an arbitrary parameter $\theta \in \Theta$ and the true parameter that generated the data, which we will call θ_0
- We'll show that the MLE is asymptotically efficient, under certain regularity conditions
- Under what? REGULARITY CONDITIONS !

Regularity Conditions

- Recall how the Cramér-Rao Lower Bound required some conditions:

$$\mathrm{Var}_\theta(T(\vec{X})) < \infty \quad \forall \theta \in \Theta \qquad \text{and} \qquad \frac{\partial}{\partial \theta} \mathbb{E}_\theta[T(\vec{X})] = \int_{\mathcal{X}} \frac{\partial}{\partial \theta} [T(\vec{x}) \cdot f_\theta(\vec{x})] \, d\vec{x}$$

- Such conditions are generically referred to as *regularity conditions*, and they're used to rule out various pathological counterexamples and edge cases

- The exact regularity conditions for our next result are quite technical and not worth getting involved with in this course

- Instead, we will go with four *sufficient* regularity conditions that are relatively easy to check, and which are satisfied by many common parametric models

Poll Time!

The MLE is Often Asymptotically Normal

- Theorem 5.9: Let $X_1, X_2, \ldots \overset{iid}{\sim} f_{\theta_0}$, and let $\hat{\theta}_n(\mathbf{X}_n)$ be the MLE of θ_0 based on a sample of size n. Suppose the following regularity conditions hold:

  - Θ is an open interval (not necessarily finite) in $\mathbb{R}$
  - The log-likelihood $\ell(\theta \mid \mathbf{x}_n)$ is three times continuously differentiable in θ
  - The support of f_θ does not depend on θ
  - $I_1(\theta) < \infty$ for all $\theta \in \Theta$

  Then

$$\sqrt{n}[\hat{\theta}_n(\mathbf{X}_n) - \theta_0] \xrightarrow{d} \mathcal{N}\left(0, \frac{1}{I_1(\theta_0)}\right).$$

  That is, $\hat{\theta}_n(\mathbf{X}_n)$ is a consistent and asymptotically efficient estimator of θ_0.

Proof (sketch). Take a Taylor series of $\ell'(\hat{\theta}_n)$ around θ_0. When n is large,

$$\ell'(\hat{\theta}_n) \approx \ell'(\theta_0) + (\hat{\theta}_n - \theta_0)\cdot \ell''(\theta_0)$$

Trust me on this.

$$\Rightarrow 0 \approx \ell'(\theta_0) + (\hat{\theta}_n - \theta_0)\cdot \ell''(\theta_0)$$

$$\Rightarrow (\hat{\theta}_n - \theta_0) \approx -\frac{\ell'(\theta_0)}{\ell''(\theta_0)}$$

$$\Rightarrow \sqrt{n}(\hat{\theta}_n - \theta_0) \approx \frac{-\frac{1}{\sqrt{n}} l'(\theta_0)}{\frac{1}{n} l''(\theta_0)} \quad \begin{matrix}① \\ ②\end{matrix}$$

① : $-\frac{1}{\sqrt{n}} l'(\theta_0) = -\frac{1}{\sqrt{n}} S(\theta_0 | \vec{X}_n)$

$= -\frac{1}{\sqrt{n}} \sum_{i=1}^{n} S(\theta_0 | X_i)$

$= \sqrt{n}\left(-\frac{1}{n} \sum_{i=1}^{n} S(\theta_0 | X_i) - 0 \right)$

$= \sqrt{n}\left(\overline{-S(\theta_0 | X)}_n - \underbrace{\mathbb{E}_{\theta_0}[S(\theta_0 | X_i)]}_{=0} \right)$

$\xrightarrow{d} N\left(0, \underbrace{\mathrm{Var}_{\theta_0}(-S(\theta_0 | X_i))}_{= I_1(\theta_0)}\right)$ by the CLT

$= N(0, I_1(\theta_0))$.

② : $\frac{1}{n} l''(\theta_0) = \frac{1}{n} \frac{\partial}{\partial \theta} S(\theta | \vec{X}_n)\Big|_{\theta=\theta_0}$

$= \frac{1}{n} \sum_{i=1}^{n} \frac{\partial}{\partial \theta} S(\theta | X_i)\Big|_{\theta=\theta_0}$

$= \overline{\left(\frac{\partial}{\partial \theta} S(\theta | X)\Big|_{\theta=\theta_0} \right)}_n$

$\xrightarrow{p} \mathbb{E}_{\theta_0}\left[\frac{\partial}{\partial \theta} S(\theta | X_i)\Big|_{\theta=\theta_0} \right]$ by the WLLN

$= -I_1(\theta_0)$ by Theorem 2.11

By Slutsky's theorem, $\sqrt{n}(\hat{\theta}_n - \theta_0) \approx \frac{-\frac{1}{\sqrt{n}} l'(\theta_0)}{\frac{1}{n} l''(\theta_0)} \xrightarrow{d} \frac{1}{-I_1(\theta_0)} \cdot N(0, I_1(\theta_0)) = N\left(0, \frac{1}{I_1(\theta_0)}\right)$.

So $\hat{\theta}_n$ is asymptotically efficient!

Consistency follows immediately by Slutsky. "$\square$"

A Useful Corollary

- Theorem 5.10: Suppose the hypotheses of Theorem 5.9 hold, and that $\tau : \Theta \to \mathbb{R}$ is continuously differentiable with $\tau'(\theta_0) \neq 0$. Then

$$\sqrt{n}[\tau(\hat{\theta}_n(\mathbf{X}_n)) - \tau(\theta_0)] \xrightarrow{d} \mathcal{N}\left(0, \frac{[\tau'(\theta_0)]^2}{I_1(\theta_0)}\right).$$

That is, $\tau(\hat{\theta}_n(\mathbf{X}_n))$ is a consistent and asymptotically efficient estimator of $\tau(\theta_0)$.

Proof: Exercise! □

Asymptotically Efficient MLEs: Examples

- Example 5.20: Let $X_1, X_2, \ldots, X_n \overset{iid}{\sim} \mathcal{N}(\mu, \sigma^2)$, where $\mu \in \mathbb{R}$ and σ^2 is known. Find the asymptotic distribution of the MLE of μ.

Check conditions of Theorem 5.9:

① $\Theta = \mathbb{R}$ open in $\mathbb{R}$ ✓

② $\ell'''(\mu|x)$ cts in μ ✓

③ Support of $N(\mu, \sigma^2)$ is $\mathbb{R}$, free of μ ✓

④ $I_1(\mu) = 1/\sigma^2 < \infty \quad \forall \mu \in \mathbb{R}$.

$\ell(\mu|x) = c - \frac{(x-\mu)^2}{2\sigma^2}$, where c is free of μ

$\ell'(\mu|x) = \frac{x-\mu}{\sigma^2}$

$\ell''(\mu|x) = -\frac{1}{\sigma^2} = S'(\mu|x) \Rightarrow I_2(\mu) = -\mathbb{E}_\mu[-\frac{1}{\sigma^2}] = 1/\sigma^2$

$\ell'''(\mu|x) = 0$ cts in μ

So by Theorem 5.9, $\hat{\mu} = \bar{X}_n$ is asymptotically efficient, with asymptotic distribution $N(0, \sigma^2)$.

OR: Since $X_i \overset{iid}{\sim} N(\mu, \sigma^2)$, $\sqrt{n}(\bar{X}_n - \mu) \xrightarrow{d} N(0, \sigma^2)$

$\sigma^2 = 1/I_1(\mu)$.

So asymptotically efficient.

Asymptotically Efficient MLEs: Examples

- Example 5.21: Let $X_1, X_2, \ldots, X_n \overset{iid}{\sim}$ Bernoulli (p), where $p \in (0,1)$. Find the asymptotic distribution of the MLE of p, and then that of $1/p$.

Check conditions:

① $\mathcal{H} = (0,1)$ open in $\mathbb{R}$ ✓

② $\ell'''(p|x)$ continuous for $p \in (0,1)$ ✓

③ Support of Bernoulli(p) is $\{0,1\}$ free of p ✓

④ $I_1(p) = \frac{1}{p(1-p)} < \infty \quad \forall\, p \in (0,1)$

$L(p|x) = p^x(1-p)^{1-x}$

$\ell(p|x) = x \cdot \log(p) + (1-x) \cdot \log(1-p)$

$\ell'''(p|x) = \frac{2x}{p^3} - \frac{2(1-x)}{(1-p)^3}$

$S(p|x) = \frac{x}{p} - \frac{1-x}{1-p}$

$S'(p|x) = -\frac{x}{p^2} + \frac{(1-x)}{(1-p)^2}$

$I_1(p) = -\mathbb{E}_p[S'(p|X)] = \frac{1}{p(1-p)}$

By Theorem 5.9, $\sqrt{n}(\bar{X}_n - p) \xrightarrow{d} N\left(0, \frac{1}{I_1(p)}\right)$

$= N(0, p(1-p))$

Let $g(p) = 1/p$. Then $g'(p) = -1/p^2$, cts on $(0,1)$.

So by Theorem 5.10, $\sqrt{n}\left(\frac{1}{\bar{X}_n} - \frac{1}{p}\right) \xrightarrow{d} N\left(0, \frac{1}{p^4} p(1-p)\right)$

$= N\left(0, \frac{1-p}{p^3}\right)$

The MLE Isn't Always Asymptotically Normal

- Example 5.22: Let $X_1, X_2, \ldots, X_n \overset{iid}{\sim} \text{Unif}(0, \theta)$, where $\theta > 0$. Show that the MLE of θ is not asymptotically normal.

$\hat{\theta}_{MLE} = X_{(n)}$. If $\sqrt{n}(X_{(n)} - \theta) \xrightarrow{d} N(0, ?)$, then $Y_n := \sqrt{n}(\theta - X_{(n)}) \xrightarrow{d} N(0, ?)$ too.

But $P_\theta(Y_n \le y)$

$= P_\theta(\theta - X_{(n)} \le y/\sqrt{n})$

$= P_\theta(-X_{(n)} \le y/\sqrt{n} - \theta)$

$= P_\theta(X_{(n)} \ge \theta - y/\sqrt{n})$

$= 1 - P_\theta(X_{(n)} \le \theta - y/\sqrt{n})$

$= 1 - \left(\frac{\theta - y/\sqrt{n}}{\theta}\right)^n$

$= 1 - (1 - y/\sqrt{n})^n$

$\xrightarrow{n \to \infty} \begin{cases} 1, & y \ge 0 \\ 0, & y < 0 \end{cases}$ (degenerate at 0)

If n instead of $\sqrt{n}$,

$1 - (1 - y/n)^n \xrightarrow{n \to \infty} 1 - e^{-y} \sim \text{Exp}(1)$.

Approximate Tests and Intervals

- We've seen that a lot of statistics are asymptotically normal
- What about test statistics?
- If we're willing to approximate a test statistic (whose exact distribution we might not know for fixed n) by one with a Normal distribution, we can perform tests and create intervals that we couldn't have before
- As in Modules 3 and 4, we'll start off with tests and then use the test statistics from those to construct confidence intervals

Wilks' Theorem

- Recall the LRT statistic for testing $H_0 : \theta = \theta_0$ versus $H_A : \theta \neq \theta_0$ was given by $\lambda(\mathbf{X}_n) = \frac{L(\theta_0|\mathbf{X}_n)}{L(\hat{\theta}|\mathbf{X}_n)}$, where $\hat{\theta} = \hat{\theta}(\mathbf{X}_n)$ is the unrestricted MLE of θ based on $\mathbf{X}_n$

- Amazingly, the LRT statistic always converges in distribution to a known distribution, regardless of the statistical model (assuming it's nice enough)

- Theorem 5.11 (**Wilks' theorem**): Let $X_1, X_2, \ldots \overset{iid}{\sim} f_\theta$, where the model satisfies the same regularity conditions as in Theorem 5.9. If we test $H_0 : \theta = \theta_0$ versus $H_A : \theta \neq \theta_0$ using $\lambda(\mathbf{X}_n)$, then

$$-2\log\left(\lambda(\mathbf{X}_n)\right) \xrightarrow{d} \chi^2_{(1)}. \quad \text{under } H_0.$$

Poll Time!

$$\lambda(\vec{X}) = \frac{\sup_{\theta \in \Theta_0} L(\theta | \vec{X})}{\sup_{\theta \in \Theta} L(\theta | \vec{X})} \in (0,1)$$

$$\Rightarrow \log(\lambda(\vec{X})) \in (-\infty, 0)$$

$$\Rightarrow -2\log(\lambda(\vec{X})) \in (0, \infty)$$ always positive!

Approximate LRTs: Examples

- Example 5.23: Let $X_1, X_2, \ldots, X_n \overset{iid}{\sim}$ Bernoulli (p), where $p \in (0,1)$. Construct an approximate size-α LRT of $H_0 : p = p_0$ versus $H_A : p \neq p_0$.

Example 3.23: $\lambda(\vec{X}_n) = \left(\frac{p_0}{\bar{X}_n}\right)^{n\bar{X}_n} \left(\frac{1-p_0}{1-\bar{X}_n}\right)^{n-n\bar{X}_n}$

$\Rightarrow \log(\lambda(\vec{X}_n)) = n\left[\bar{X}_n \log\left(p_0/\bar{X}_n\right) + (1-\bar{X}_n)\cdot\log\left(\frac{1-p_0}{1-\bar{X}_n}\right)\right]$

$\Rightarrow -2\cdot\log(\lambda(\vec{X}_n)) = -2n\left[\bar{X}_n \log\left(p_0/\bar{X}_n\right) + (1-\bar{X}_n)\cdot\log\left(\frac{1-p_0}{1-\bar{X}_n}\right)\right]$

By Wilks' theorem, $R = \left\{\vec{x} \in \mathcal{X}^n : -2n\left(\bar{x}\cdot\log(p_0/\bar{x}) + (1-\bar{x})\cdot\log\left(\frac{1-p_0}{1-\bar{x}}\right)\right) \geq \chi^2_{(1),1-\alpha}\right\}$

is the rejection region of an approximate size-α test of $H_0: p = p_0$ vs $H_A: p \neq p_0$.

Approximate LRTs: Examples

- Example 5.24: Let $X_1, X_2, \ldots, X_n \overset{iid}{\sim} \mathcal{N}\left(\mu, \sigma^2\right)$, where $\mu \in \mathbb{R}$. Construct an approximate size-α LRT of $H_0 : \mu = \mu_0$ versus $H_A : \mu \neq \mu_0$.

Example 3.21: $\lambda(\vec{x}_n) = \exp\left(\frac{-n}{2\sigma^2}(\bar{x}_n - \mu_0)^2\right)$

$\Rightarrow \log(\lambda(\vec{x}_n)) = \frac{-n}{2\sigma^2}(\bar{x}_n - \mu_0)^2$

$\Rightarrow -2 \cdot \log(\lambda(\vec{x}_n)) = \frac{n}{\sigma^2}(\bar{x}_n - \mu_0)^2$

By Wilks' theorem, $R = \{\vec{x} \in \mathcal{X}^n : \frac{n}{\sigma^2}(\bar{x} - \mu_0)^2 \geq \chi^2_{(1,1-\alpha)}\}$

Wald Tests

- Definition 5.7: Let $X_1, X_2, \ldots, X_n \overset{iid}{\sim} f_\theta$. For testing $H_0 : \theta = \theta_0$ versus $H_A : \theta \neq \theta_0$, a **Wald test** is a test based on the **Wald statistic**

$$W_n(\mathbf{X}_n) = (\hat{\theta} - \theta_0)^2 I_n(\hat{\theta}),$$

"plug-in Fisher information"

where $\hat{\theta} = \hat{\theta}(\mathbf{X}_n)$ is the (unrestricted) MLE.

- Theorem 5.12: Let $X_1, X_2, \ldots, X_n \overset{iid}{\sim} f_\theta$, where the model satisfies the same regularity conditions as in Theorem 5.9. If we test $H_0 : \theta = \theta_0$ versus $H_A : \theta \neq \theta_0$ using $W_n(\mathbf{X}_n)$, then

$$W_n(\mathbf{X}_n) \xrightarrow{d} \chi^2_{(1)}.$$ under H_0. Proof: Exercise. □

Because $(\hat{\theta} - \theta_0)\sqrt{I_n(\hat{\theta})} \xrightarrow{d} N(0,1)$ under H_0 (this is usually easier to work with)

Wald Tests: Examples

- Example 5.25: Let $X_1, X_2, \ldots, X_n \overset{iid}{\sim} \text{Bernoulli}(p)$, where $p \in (0,1)$. Construct an approximate size-α Wald test of $H_0 : p = p_0$ versus $H_A : p \neq p_0$.

$W_n(\vec{X}_n) = (\hat{p} - p_0)^2 I_n(\hat{p})$, where $\hat{p} = \bar{X}_n$.

Fisher information: $I_n(p) = \frac{n}{p(1-p)}$.

$\Rightarrow I_n(\hat{p}) = \frac{n}{\bar{X}_n(1-\bar{X}_n)}$

So the Wald statistic is $W_n(\vec{X}_n) = \frac{(\bar{X}_n - p_0)^2 \cdot n}{\bar{X}_n(1-\bar{X}_n)}$

So by Theorem 5.12, $R = \left\{\vec{x} \in \mathcal{X}^n : \frac{(\bar{x} - p_0)^2 \cdot n}{\bar{x}(1-\bar{x})} > \chi^2_{(0,1-\alpha)}\right\}$ is the rejection region of an approximate size-α test of $H_0 : p = p_0$ vs $H_A : p \neq p_0$.

OR: By Theorem 5.12, $R' = \left\{\vec{x} \in \mathcal{X}^n : \left| \frac{(\bar{x} - p_0)}{\sqrt{\bar{x}(1-\bar{x})/n}} \right| > z_{1-\alpha/2}\right\}$ is the rejection region....

Wald Tests: Examples

- Example 5.26: Let $X_1, X_2, \ldots, X_n \overset{iid}{\sim} \mathcal{N}(\mu, \sigma^2)$, where $\mu \in \mathbb{R}$. Construct an approximate size-α Wald test of $H_0 : \mu = \mu_0$ versus $H_A : \mu \neq \mu_0$.

$I_n(\mu) = n/\sigma^2$ (Example 5.20, etc), $\hat{\mu}_{MLE} = \bar{X}_n$.

So $W_n(\vec{X}_n) = \dfrac{(\bar{X}_n - \mu_0)^2 \cdot n}{\sigma^2} = \left(\dfrac{\bar{X}_n - \mu_0}{\sqrt{\sigma^2/n}}\right)^2$.

By Theorem 5.12, $R = \{\vec{x} \in \mathcal{X}^n : \frac{(\bar{x}-\mu_0)^2 \cdot n}{\sigma^2} \geq \chi^2_{(0,1-\alpha)}\}$ is the rejection region of an approximate size-α test of $H_0: \mu = \mu_0$ vs $H_A: \mu \neq \mu_0$.

In this case, it's exact!

OR: By Theorem 5.12, $R' = \left\{\vec{x} \in \mathcal{X}^n : \left|\dfrac{\bar{X}_n - \mu_0}{\sqrt{\sigma^2/n}}\right| \geq z_{1-\alpha/2}\right\}$ is the rejection region.

Hey, it's our two-sided z-test!

Score Tests

- Definition 5.8: Let $X_1, X_2, \ldots, X_n \overset{iid}{\sim} f_\theta$. For testing $H_0 : \theta \in \Theta_0$ versus $H_A : \theta \in \Theta_0^c$, a **score test** (also called a **Rao test** or a **Lagrange multiplier test** is a test based on the **score statistic**

$$R_n(\mathbf{X}_n) = \frac{[S_n(\hat{\theta}_0 \mid \mathbf{X}_n)]^2}{I_n(\hat{\theta}_0)},$$

where $\hat{\theta}_0 = \hat{\theta}_0(\mathbf{X}_n) = \underset{\theta \in \Theta_0}{\operatorname{argmax}} L(\theta \mid \mathbf{X}_n)$ is the restricted MLE under H_0.

- Theorem 5.13: Let $X_1, X_2, \ldots, X_n \overset{iid}{\sim} f_\theta$, where the model satisfies the same regularity conditions as in Theorem 5.9. If we test $H_0 : \theta \in \Theta_0$ versus $H_A : \theta \in \Theta_0^c$ using $R_n(\mathbf{X}_n)$, then

$$R_n(\mathbf{X}_n) \xrightarrow{d} \chi^2_{(1)}. \quad \text{under } H_0.$$

Because $\dfrac{S_n(\hat{\theta}_0 | \vec{X}_n)}{\sqrt{I_n(\hat{\theta}_0)}} \xrightarrow{d} N(0,1)$ under H_0 (this is usually easier to work with)

Score Tests: Examples

$\Theta_0 = \{p_0\}$

- Example 5.27: Let $X_1, X_2, \ldots, X_n \overset{iid}{\sim}$ Bernoulli (p), where $p \in (0,1)$. Construct a (approximate) size-α score test of $H_0 : p = p_0$ versus $H_A : p \neq p_0$.

$$R_n(\vec{X}_n) = \frac{[S_n(\hat{p}_0 | \vec{X}_n)]^2}{I_n(\hat{p}_0)} = \frac{[S_n(p_0 | \vec{X}_n)]^2}{I_n(p_0)}$$

$$= n^2 \left[\frac{\bar{X}_n}{p_0} - \frac{1-\bar{X}_n}{1-p_0} \right]^2 \frac{p_0(1-p_0)}{n}$$

$$= \frac{(\bar{X}_n - p_0)^2}{p_0(1-p_0)/n}$$ check!

$L(p|\vec{x}) = p^{\sum x_i}(1-p)^{n-\sum x_i}$

$\ell(p|\vec{x}) = \sum x_i \log(p) + (n - \sum x_i)\log(1-p)$

$S(p|\vec{x}) = \frac{\sum x_i}{p} - \frac{n - \sum x_i}{1-p} = n\left[\frac{\bar{x}_n}{p} - \frac{1-\bar{x}_n}{1-p}\right]$

$S'(p|\vec{x}) = -\frac{\sum x_i}{p^2} + \frac{n - \sum x_i}{(1-p)^2}$

$I_n(p) = E_p\left[-\frac{\sum x_i}{p} + \frac{n - \sum x_i}{(1-p)^2}\right] = \frac{n}{p(1-p)}$

By Theorem 5.13, $R = \left\{ \vec{x} \in \mathcal{X}^n : \frac{(\bar{x} - p_0)^2}{p_0(1-p_0)/n} \geq \chi^2_{(1),1-\alpha} \right\}$ is the rejection of an approximate size-α test of $H_0 : p = p_0$ vs $H_A : p \neq p_0$.

OR: By Theorem 5.13, $R' = \left\{ \vec{x} \in \mathcal{X}^n : \left| \frac{\bar{x} - p_0}{\sqrt{p_0(1-p_0)/n}} \right| \geq z_{1-\alpha/2} \right\}$ is the rejection region of...

Score Tests: Examples

- Example 5.28: Let $X_1, X_2, \ldots, X_n \overset{iid}{\sim} \mathcal{N}(\mu, \sigma^2)$, where $\mu \in \mathbb{R}$. Construct an approximate size-α score test of $H_0 : \mu = \mu_0$ versus $H_A : \mu \neq \mu_0$.

$S(\mu|\vec{x}) = \frac{n(\bar{x}-\mu)}{\sigma^2}$ so $R_n(\vec{x}_n) = \frac{[S(\mu_0|\vec{x}_n)]^2}{I_n(\mu_0)}$

$\Rightarrow S'(\mu|\vec{x}) = -\frac{n}{\sigma^2}$

$= \frac{n^2(\bar{x}_n-\mu_0)^2}{\sigma^4} \cdot \frac{\sigma^2}{n}$

$\Rightarrow I_n(\mu) = n/\sigma^2$

$= \left(\frac{\bar{x}_n-\mu_0}{\sqrt{\sigma^2/n}}\right)^2$

By Theorem 5.13, $R = \{\vec{x} \in \mathcal{X}^n : \frac{(\bar{x}-\mu_0)^2 \cdot n}{\sigma^2} \geq \chi^2_{(0,1-\alpha)}\}$ is the rejection region of an approximate size-α test of $H_0: \mu = \mu_0$ vs. $H_A: \mu \neq \mu_0$.

↑ In this case, it's exact!

OR: By Theorem 5.13, $R' = \{\vec{x} \in \mathcal{X}^n : \left|\frac{\bar{x}-\mu_0}{\sqrt{\sigma^2/n}}\right| \geq z_{1-\alpha/2}\}$ is the rejection region of...

Hey, it's our two-sided Z-test again!

The Trinity of Tests

- The LRT, the Wald test, and the score test form the backbone of classical hypothesis testing
- Observe that under H_0, all three tests are asymptotically equivalent (i.e., all three test statistics all converge in distribution to a $\chi^2_{(1)}$)
- For this reason, the three tests are sometimes collectively referred to as the **trinity of tests**
- Although asymptotically equivalent, the speed of convergence to $\chi^2_{(1)}$ can be quite different for each one – for small n, they can be quite different in terms of power and other "small-sample" properties
- One might tell you to reject H_0 while another might not!

Fact: $\ell(\theta|x) = a\theta^2 + b\theta + c$ for some $a, b, c \in \mathbb{R}$, then all three give the same result (1982)

Approximate Confidence Intervals

- Using any of the asymptotic tests to test $H_0 : \theta = \theta_0$ versus $H_A : \theta \neq \theta_0$, it's sometimes possible to invert any of the test statistics to obtain an approximate $(1 - \alpha)$-confidence interval for θ
- Out of the three, the LRT is usually the hardest to invert into an actual interval, and the Wald statistic is usually the easiest
- In practice, you can always try to use numerical solvers when the algebra doesn't work
- For Wald and score intervals, the standard recipe is to take the square root of the test statistic and compare it to $\mathcal{N}(0, 1)$

Approximate Confidence Intervals: Examples

- Example 5.29: Let $X_1, X_2, \ldots, X_n \overset{iid}{\sim} \text{Bernoulli}(p)$, where $p \in (0,1)$. Construct a 95% confidence interval for p based on the Wald statistic.

an approximate

Example 5.25: $P_p\left(\frac{|\bar{X}_n - p|}{\sqrt{\bar{X}_n(1-\bar{X}_n)/n}} < z_{1-\alpha/2} \right) \approx 1-\alpha$

$= P_p\left(-z_{1-\alpha/2} < \frac{p - \bar{X}_n}{\sqrt{\bar{X}_n(1-\bar{X}_n)/n}} < z_{1-\alpha/2} \right)$

$= P_p\left(\bar{X}_n - z_{1-\alpha/2} \cdot \sqrt{\frac{\bar{X}_n(1-\bar{X}_n)}{n}} < p < \bar{X}_n + z_{1-\alpha/2} \cdot \sqrt{\frac{\bar{X}_n(1-\bar{X}_n)}{n}} \right)$

$\Rightarrow \left(\bar{X}_n - z_{1-\alpha/2}\sqrt{\frac{\bar{X}_n(1-\bar{X}_n)}{n}},\ \bar{X}_n + z_{1-\alpha/2}\sqrt{\frac{\bar{X}_n(1-\bar{X}_n)}{n}} \right)$ is an approximate $(1-\alpha)$-CI for p

- This confidence interval shows up everywhere in polling (and is a staple of introductory Statistics classes); its half-length is called the **margin of error**

Almost always with $\alpha = 0.05 \Rightarrow z_{1-\alpha/2} \approx 1.96$

Approximate Confidence Intervals: Examples

- Example 5.30: Let $X_1, X_2, \ldots, X_n \overset{iid}{\sim} \text{Bernoulli}(p)$, where $p \in (0,1)$. Construct an approximate 95% confidence interval for $\log\left(\frac{p}{1-p}\right)$ based on the Wald statistic.

From Example 5.29,

$$1-\alpha \approx P_p\left(\bar{X}_n - z_{1-\alpha/2}\sqrt{\frac{\bar{X}_n(1-\bar{X}_n)}{n}} < p < \bar{X}_n + z_{1-\alpha/2}\sqrt{\frac{\bar{X}_n(1-\bar{X}_n)}{n}}\right)$$

$$= P_p\left(\log\left(\frac{\bar{X}_n - z_{1-\alpha/2}\sqrt{\frac{\bar{X}_n(1-\bar{X}_n)}{n}}}{1-\bar{X}_n + z_{1-\alpha/2}\sqrt{\frac{\bar{X}_n(1-\bar{X}_n)}{n}}}\right) < \log\left(\frac{p}{1-p}\right) < \log\left(\frac{\bar{X}_n + z_{1-\alpha/2}\sqrt{\frac{\bar{X}_n(1-\bar{X}_n)}{n}}}{1-\bar{X}_n - z_{1-\alpha/2}\sqrt{\frac{\bar{X}_n(1-\bar{X}_n)}{n}}}\right)\right)$$

So $\left(\log\left(\frac{\bar{X}_n - z_{1-\alpha/2}\sqrt{\frac{\bar{X}_n(1-\bar{X}_n)}{n}}}{1-\bar{X}_n + z_{1-\alpha/2}\sqrt{\frac{\bar{X}_n(1-\bar{X}_n)}{n}}}\right), \log\left(\frac{\bar{X}_n + z_{1-\alpha/2}\sqrt{\frac{\bar{X}_n(1-\bar{X}_n)}{n}}}{1-\bar{X}_n - z_{1-\alpha/2}\sqrt{\frac{\bar{X}_n(1-\bar{X}_n)}{n}}}\right)\right)$ is an approximate $(1-\alpha)$-CI for $\log\left(\frac{p}{1-p}\right)$.

Approximate Confidence Intervals: Examples

- Example 5.31: Let $X_1, X_2, \ldots, X_n \overset{iid}{\sim} \text{Poisson}(\lambda)$, where $\lambda > 0$. Construct an approximate 95% confidence interval for λ based on the Wald statistic.

MLE: $\hat{\lambda} = \bar{X}_n$.

$\ell(\lambda|\vec{x}) = -n\lambda + \sum x_i \cdot \log(\lambda) + c$, where c is free of λ

$\Rightarrow S(\lambda|\vec{x}) = -n + \frac{\sum x_i}{\lambda}$

$\Rightarrow S'(\lambda|\vec{x}) = -\frac{\sum x_i}{\lambda^2}$

$\Rightarrow I_n(\lambda) = -E_\lambda\left[-\frac{\sum x_i}{\lambda^2}\right] = \frac{n}{\lambda}$

$\Rightarrow I_n(\hat{\lambda}) = \frac{n}{\bar{X}_n}$

So $W_n(\vec{X}_n) = \frac{(\bar{X}_n - \lambda)^2 \cdot n}{\bar{X}_n} = \frac{(\bar{X}_n - \lambda)^2}{\bar{X}_n/n}$

So $1-\alpha \approx P_\lambda\left(-z_{1-\alpha/2} < \frac{\lambda - \bar{X}_n}{\sqrt{\bar{X}_n/n}} < z_{1-\alpha/2}\right)$

$= P_\lambda\left(\bar{X}_n - z_{1-\alpha/2}\sqrt{\frac{\bar{X}_n}{n}} < \lambda < \bar{X}_n + z_{1-\alpha/2}\sqrt{\frac{\bar{X}_n}{n}}\right)$

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

is an approximate $(1-\alpha)$-CI for λ.

When the Fisher Information Causes Problems...

- When f_θ is too complicated to allow for exact $(1-\alpha)$-confidence intervals, it's standard practice to use Wald intervals and score intervals
- But there might be another problem: calculating the Fisher information I.C.'s.
- In real-life multiparameter models, $I_n(\boldsymbol{\theta})$ is a matrix and is often impossible to work out directly, which makes calculating $I_n(\hat{\boldsymbol{\theta}}_0)$ or $I_n(\hat{\boldsymbol{\theta}})$ futile
- When this happens, people like to swap $I_n(\cdot)$ with $J_n(\cdot)$ in the Wald and score statistics ... but is this justified???
- Yes — it can be shown $J_n(\hat{\Theta}_n)$ is a consistent estimator of $I_n(\Theta_0)$
- Moreover, in a famous 1978 paper, Efron and Hinkley showed empirically that $J_n(\hat{\theta})$ is *superior* to $I_n(\hat{\theta})$ Optional reading, if you're curious