

Econ 722 – Advanced Econometrics IV

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Lecture #1 – Decision Theory

Statistical Decision Theory

The James-Stein Estimator

Decision Theoretic Preliminaries

Parameter $\theta \in \Theta$

Unknown state of nature, from parameter space Θ

Observed Data

Observe X with distribution F_θ from a sample space \mathcal{X}

Estimator $\hat{\theta}$

An estimator (aka a decision rule) is a function from \mathcal{X} to Θ

Loss Function $L(\theta, \hat{\theta})$

A function from $\Theta \times \Theta$ to \mathbb{R} that gives the cost we incur if we report $\hat{\theta}$ when the true state of nature is θ .

Examples of Loss Functions

$$L(\theta, \hat{\theta}) = (\theta - \hat{\theta})^2$$

squared error loss

$$L(\theta, \hat{\theta}) = |\theta - \hat{\theta}|$$

absolute error loss

$$L(\theta, \hat{\theta}) = 0 \text{ if } \theta = \hat{\theta}, 1 \text{ otherwise}$$

zero-one loss

$$L(\theta, \hat{\theta}) = \int \log \left[\frac{f(x|\theta)}{f(x|\hat{\theta})} \right] f(x|\theta) dx$$

Kullback–Leibler loss

(Frequentist) Risk of an Estimator $\hat{\theta}$

$$R(\theta, \hat{\theta}) = \mathbb{E}_{\theta} [L(\theta, \hat{\theta})] = \int L(\theta, \hat{\theta}(x)) dF_{\theta}(x)$$

The frequentist decision theorist seeks to evaluate, for each θ , how much he would “expect” to lose if he used $\hat{\theta}(X)$ repeatedly with varying X in the problem.

(Berger, 1985)

Example: Squared Error Loss

$$R(\theta, \hat{\theta}) = \mathbb{E}_{\theta} [(\theta - \hat{\theta})^2] = \text{MSE} = \text{Var}(\hat{\theta}) + \text{Bias}_{\theta}^2(\hat{\theta})$$

Bayes Risk and Maximum Risk

Comparing Risk

$R(\theta, \hat{\theta})$ is a *function* of θ rather than a single number. We want an estimator with low risk, but how can we compare?

Maximum Risk

$$\bar{R}(\hat{\theta}) = \sup_{\theta \in \Theta} R(\theta, \hat{\theta})$$

Bayes Risk

$$r(\pi, \hat{\theta}) = \mathbb{E}_{\pi} \left[R(\theta, \hat{\theta}) \right], \text{ where } \pi \text{ is a prior for } \theta$$

Bayes and Minimax Rules

Minimize the Maximum or Bayes risk over all estimators $\tilde{\theta}$

Minimax Rule/Estimator

$\hat{\theta}$ is **minimax** if

$$\sup_{\theta \in \Theta} R(\theta, \hat{\theta}) = \inf_{\tilde{\theta}} \sup_{\theta \in \Theta} R(\theta, \tilde{\theta})$$

Bayes Rule/Estimator

$\hat{\theta}$ is a **Bayes rule** with respect to prior π if

$$r(\pi, \hat{\theta}) = \inf_{\tilde{\theta}} r(\pi, \tilde{\theta})$$

Recall: Bayes' Theorem and Marginal Likelihood

Let π be a prior for θ . By Bayes' theorem, the **posterior** $\pi(\theta|\mathbf{x})$ is

$$\pi(\theta|\mathbf{x}) = \frac{f(\mathbf{x}|\theta)\pi(\theta)}{m(\mathbf{x})}$$

where the **marginal likelihood** $m(\mathbf{x})$ is given by

$$m(\mathbf{x}) = \int f(\mathbf{x}|\theta)\pi(\theta) d\theta$$

Posterior Expected Loss

Posterior Expected Loss

$$\rho(\pi(\theta|\mathbf{x}), \hat{\theta}) = \int L(\theta, \hat{\theta}) \pi(\theta|\mathbf{x}) d\theta$$

Bayesian Decision Theory

Choose an estimator that minimizes posterior expected loss.

Easier Calculation

Since $m(\mathbf{x})$ does not depend on θ , to minimize $\rho(\pi(\theta|\mathbf{x}), \hat{\theta})$ it suffices to minimize $\int L(\theta, \hat{\theta}) f(\mathbf{x}|\theta) \pi(\theta) d\theta$.

Question

Is there a relationship between Bayes risk, $r(\pi, \hat{\theta}) \equiv \mathbb{E}_{\pi}[R(\theta, \hat{\theta})]$, and posterior expected loss?

Bayes Risk vs. Posterior Expected Loss

Theorem

$$r(\pi, \hat{\theta}) = \int \rho(\pi(\theta|\mathbf{x}), \hat{\theta}(\mathbf{x})) m(\mathbf{x}) d\mathbf{x}$$

Proof

$$\begin{aligned} r(\pi, \hat{\theta}) &= \int R(\theta, \hat{\theta}) \pi(\theta) d\theta = \int \left[\int L(\theta, \hat{\theta}(\mathbf{x})) f(\mathbf{x}|\theta) d\mathbf{x} \right] \pi(\theta) d\theta \\ &= \int \int L(\theta, \hat{\theta}(\mathbf{x})) [f(\mathbf{x}|\theta) \pi(\theta)] d\mathbf{x} d\theta \\ &= \int \int L(\theta, \hat{\theta}(\mathbf{x})) [\pi(\theta|\mathbf{x}) m(\mathbf{x})] d\mathbf{x} d\theta \\ &= \int \left[\int L(\theta, \hat{\theta}(\mathbf{x})) \pi(\theta|\mathbf{x}) d\theta \right] m(\mathbf{x}) d\mathbf{x} \\ &= \int \rho(\pi(\theta|\mathbf{x}), \hat{\theta}(\mathbf{x})) m(\mathbf{x}) d\mathbf{x} \end{aligned}$$

Finding a Bayes Estimator

Hard Problem

Find the **function** $\hat{\theta}(\mathbf{x})$ that minimizes $r(\pi, \hat{\theta})$.

Easy Problem

Find the **number** $\hat{\theta}$ that minimizes $\rho(\pi(\theta|\mathbf{x}), \hat{\theta})$

Punchline

Since $r(\pi, \hat{\theta}) = \int \rho(\pi(\theta|\mathbf{x}), \hat{\theta}(\mathbf{x})) m(\mathbf{x}) d\mathbf{x}$, to minimize $r(\pi, \hat{\theta})$ we can set $\hat{\theta}(\mathbf{x})$ to be the value $\hat{\theta}$ that minimizes $\rho(\pi(\theta|\mathbf{x}), \hat{\theta})$.

Bayes Estimators for Common Loss Functions

Zero-one Loss

For zero-one loss, the Bayes estimator is the posterior mode.

Absolute Error Loss: $L(\theta, \hat{\theta}) = |\theta - \hat{\theta}|$

For absolute error loss, the Bayes estimator is the posterior median.

Squared Error Loss: $L(\theta, \hat{\theta}) = (\theta - \hat{\theta})^2$

For squared error loss, the Bayes estimator is the posterior mean.

Derivation of Bayes Estimator for Squared Error Loss

By definition,

$$\hat{\theta} \equiv \arg \min_{a \in \Theta} \int (\theta - a)^2 \pi(\theta | \mathbf{x}) d\theta$$

Differentiating with respect to a , we have

$$\begin{aligned} 2 \int (\theta - a) \pi(\theta | \mathbf{x}) d\theta &= 0 \\ \int \theta \pi(\theta | \mathbf{x}) d\theta &= a \end{aligned}$$

Example: Bayes Estimator for a Normal Mean

Suppose $X \sim N(\mu, 1)$ and π is a $N(a, b^2)$ prior. Then,

$$\begin{aligned}\pi(\mu|x) &\propto f(x|\mu) \times \pi(\mu) \\ &\propto \exp \left\{ -\frac{1}{2} \left[(x - \mu)^2 + \frac{1}{b^2} (\mu - a)^2 \right] \right\} \\ &\propto \exp \left\{ -\frac{1}{2} \left[\left(1 + \frac{1}{b^2} \right) \mu^2 - 2 \left(x + \frac{a}{b^2} \right) \mu \right] \right\} \\ &\propto \exp \left\{ -\frac{1}{2} \left(\frac{b^2 + 1}{b^2} \right) \left[\mu - \left(\frac{b^2 x + a}{b^2 + 1} \right) \right]^2 \right\}\end{aligned}$$

So $\pi(\mu|x)$ is $N(m, \omega^2)$ with $\omega^2 = \frac{b^2}{1+b^2}$ and $m = \omega^2 x + (1 - \omega^2)a$.

Hence the Bayes estimator for μ under squared error loss is

$$\hat{\theta}(X) = \frac{b^2 X + a}{1 + b^2}$$

Minimax Analysis

Wasserman (2004)

The advantage of using maximum risk, despite its problems, is that it does not require one to choose a prior.

Berger (1986)

Perhaps the greatest use of the minimax principle is in situations for which no prior information is available . . . but two notes of caution should be sounded. First, the minimax principle can lead to bad decision rules. . . Second, the minimax approach can be devilishly hard to implement.

Methods for Finding a Minimax Estimator

1. Direct Calculation
2. Guess a “Least Favorable” Prior
3. Search for an “Equalizer Rule”

Method 1 rarely applicable so focus on 2 and 3. . .

The Bayes Rule for a Least Favorable Prior is Minimax

Theorem

Let $\hat{\theta}$ be a Bayes rule with respect to π and suppose that for all $\theta \in \Theta$ we have $R(\theta, \hat{\theta}) \leq r(\pi, \hat{\theta})$. Then $\hat{\theta}$ is a **minimax estimator**, and π is called a **least favorable prior**.

Proof

Suppose that $\hat{\theta}$ is not minimax. Then there exists another estimator $\tilde{\theta}$ with $\sup_{\theta \in \Theta} R(\theta, \tilde{\theta}) < \sup_{\theta \in \Theta} R(\theta, \hat{\theta})$. But since

$$r(\pi, \tilde{\theta}) \equiv \mathbb{E}_{\pi} [R(\theta, \tilde{\theta})] \leq \mathbb{E}_{\pi} \left[\sup_{\theta \in \Theta} R(\theta, \tilde{\theta}) \right] = \sup_{\theta \in \Theta} R(\theta, \tilde{\theta})$$

but this implies that $\tilde{\theta}$ is *not* Bayes with respect to π since

$$r(\pi, \tilde{\theta}) \leq \sup_{\theta \in \Theta} R(\theta, \tilde{\theta}) < \sup_{\theta \in \Theta} R(\theta, \hat{\theta}) \leq r(\pi, \hat{\theta})$$

Example of Least Favorable Prior

Bounded Normal Mean

- ▶ $X \sim N(\theta, 1)$
- ▶ Squared error loss
- ▶ $\Theta = [-m, m]$ for $0 < m < 1$

Least Favorable Prior

$\pi(\theta) = 1/2$ for $\theta \in \{-m, m\}$, zero otherwise.

Resulting Bayes Rule is Minimax

$$\hat{\theta}(X) = m \tanh(mX) = m \left[\frac{\exp\{mX\} - \exp\{-mX\}}{\exp\{mX\} + \exp\{-mX\}} \right]$$

Equalizer Rules

Definition

An estimator $\hat{\theta}$ is called an **equalizer rule** if its risk function is constant: $R(\theta, \hat{\theta}) = C$ for some C .

Theorem

If $\hat{\theta}$ is an equalizer rule and is Bayes with respect to π , then $\hat{\theta}$ is **minimax** and π is **least favorable**.

Proof

$$r(\pi, \hat{\theta}) = \int R(\theta, \hat{\theta}) \pi(\theta) d\theta = \int C \pi(\theta) d\theta = C$$

Hence, $R(\theta, \hat{\theta}) \leq r(\pi, \hat{\theta})$ for all θ so we can apply the preceding theorem.

Example: $X_1, \dots, X_n \sim \text{iid Bernoulli}(p)$

Under a $\text{Beta}(\alpha, \beta)$ prior with $\alpha = \beta = \sqrt{n}/2$,

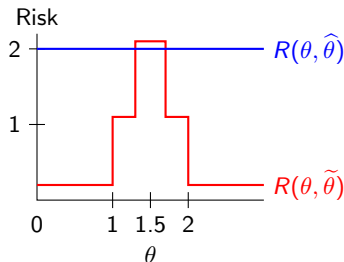
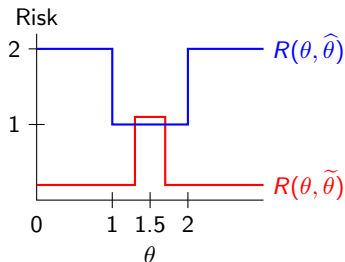
$$\hat{p}(\mathbf{x}) = \frac{n\bar{X} + \sqrt{n}/2}{n + \sqrt{n}}$$

is the Bayesian posterior mean, hence the Bayes rule under squared error loss. The risk function of \hat{p} is,

$$R(p, \hat{p}) = \frac{n}{4(n + \sqrt{n})^2}$$

which is constant in p . Hence, \hat{p} is an equalizer rule, and by the preceding theorem is minimax.

Problems with the Minimax Principle



In the left panel, $\tilde{\theta}$ is preferred by the minimax principle; in the right panel $\hat{\theta}$ is preferred. But the only difference between them is that the right panel adds an additional *fixed* loss of 1 for $1 \leq \theta \leq 2$.

Problems with the Minimax Principle

Suppose that $\Theta = \{\theta_1, \theta_2\}$, $\mathcal{A} = \{a_1, a_2\}$ and the loss function is:

	a_1	a_2
θ_1	10	10.01
θ_2	8	-8

- ▶ Minimax principle: choose a_1
- ▶ Bayes: Choose a_2 unless $\pi(\theta_1) > 0.9994$

Minimax ignores the fact that under θ_1 we can never do better than a loss of 10, and tries to prevent us from incurring a tiny additional loss of 0.01

Dominance and Admissibility

Dominance

$\hat{\theta}$ **dominates** $\tilde{\theta}$ with respect to R if $R(\theta, \hat{\theta}) \leq R(\theta, \tilde{\theta})$ for all $\theta \in \Theta$ and the inequality is strict for at least one value of θ .

Admissibility

$\hat{\theta}$ is **admissible** if no other estimator dominates it.

Inadmissibility

$\hat{\theta}$ is **inadmissible** if there is an estimator that dominates it.

Example of an Admissible Estimator

Say we want to estimate θ from $X \sim N(\theta, 1)$ under squared error loss. Is the estimator $\hat{\theta}(X) = 3$ admissible?

If not, then there is a $\tilde{\theta}$ with $R(\theta, \tilde{\theta}) \leq R(\theta, \hat{\theta})$ for all θ . Hence:

$$R(3, \tilde{\theta}) \leq R(3, \hat{\theta}) = \left\{ \mathbb{E} [\hat{\theta} - 3] \right\}^2 + \text{Var}(\hat{\theta}) = 0$$

Since R cannot be negative for squared error loss,

$$0 = R(3, \tilde{\theta}) = \left\{ \mathbb{E} [\tilde{\theta} - 3] \right\}^2 + \text{Var}(\tilde{\theta})$$

Therefore $\hat{\theta} = \tilde{\theta}$, so $\hat{\theta}$ is admissible, although very silly!

Bayes Rules are Admissible

Theorem A-1

Suppose that Θ is a discrete set and π gives strictly positive probability to each element of Θ . Then, if $\hat{\theta}$ is a Bayes rule with respect to π , it is admissible.

Theorem A-2

If a Bayes rule is unique, it is admissible.

Theorem A-3

Suppose that $R(\theta, \hat{\theta})$ is continuous in θ for all $\hat{\theta}$ and that π gives strictly positive probability to any open subset of Θ . Then if $\hat{\theta}$ is a Bayes rule with respect to π , it is admissible.

Admissible Equalizer Rules are Minimax

Theorem

Let $\hat{\theta}$ be an equalizer rule. Then if $\hat{\theta}$ is admissible, it is minimax.

Proof

Since $\hat{\theta}$ is an equalizer rule, $R(\theta, \hat{\theta}) = C$. Suppose that $\hat{\theta}$ is not minimax. Then there is a $\tilde{\theta}$ such that

$$\sup_{\theta \in \Theta} R(\theta, \tilde{\theta}) < \sup_{\theta \in \Theta} R(\theta, \hat{\theta}) = C$$

But for any θ , $R(\theta, \tilde{\theta}) \leq \sup_{\theta \in \Theta} R(\theta, \tilde{\theta})$. Thus we have shown that $\tilde{\theta}$ dominates $\hat{\theta}$, so that $\hat{\theta}$ cannot be admissible.

Minimax Implies “Nearly” Admissible

Strong Inadmissibility

We say that $\hat{\theta}$ is **strongly inadmissible** if there exists an estimator $\tilde{\theta}$ and an $\varepsilon > 0$ such that $R(\theta, \tilde{\theta}) < R(\theta, \hat{\theta}) - \varepsilon$ for all θ .

Theorem

If $\hat{\theta}$ is minimax, then it is **not** strongly inadmissible.

Example: Sample Mean, Unbounded Parameter Space

Theorem

Suppose that $X_1, \dots, X_n \sim N(\theta, 1)$ with $\Theta = \mathbb{R}$. Under squared error loss, one can show that $\hat{\theta} = \bar{X}$ is admissible.

Intuition

The proof is complicated, but effectively we view this estimator as a **limit** of a of Bayes estimator with prior $N(a, b^2)$, as $b^2 \rightarrow \infty$.

Minimaxity

Since $R(\theta, \bar{X}) = \text{Var}(\bar{X}) = 1/n$, we see that \bar{X} is an equalizer rule. Since it is admissible, it is therefore minimax.

Recall: Gauss-Markov Theorem

Linear Regression Model

$$\mathbf{y} = X\beta + \epsilon, \quad \mathbb{E}[\epsilon|X] = \mathbf{0}$$

Best Linear Unbiased Estimator

- ▶ $\text{Var}(\epsilon|X) = \sigma^2 I \Rightarrow$ then OLS has lowest variance among linear, unbiased estimators of β .
- ▶ $\text{Var}(\epsilon|X) \neq \sigma^2 I \Rightarrow$ then GLS gives a lower variance estimator.

What if we consider biased estimators and squared error loss?

Multiple Normal Means: $X \sim N(\theta, I)$

Goal

Estimate the p -vector θ using X with $L(\theta, \hat{\theta}) = \|\hat{\theta} - \theta\|^2$.

Maximum Likelihood Estimator $\hat{\theta}$

MLE = sample mean, but only one observation: $\hat{\theta} = X$.

Risk of $\hat{\theta}$

$$(\hat{\theta} - \theta)' (\hat{\theta} - \theta) = (X - \theta)' (X - \theta) = \sum_{i=1}^p (X_i - \theta_i)^2 \sim \chi_p^2$$

Since $\mathbb{E}[\chi_p^2] = p$, we have $R(\theta, \hat{\theta}) = p$.

Multiple Normal Means: $X \sim N(\theta, I)$

James-Stein Estimator

$$\hat{\theta}^{JS} = \hat{\theta} \left(1 - \frac{p-2}{\hat{\theta}'\hat{\theta}} \right) = X - \frac{(p-2)X}{X'X}$$

- ▶ Shrinks components of sample mean vector towards zero
- ▶ More elements in $\theta \Rightarrow$ more shrinkage
- ▶ MLE close to zero ($\hat{\theta}'\hat{\theta}$ small) gives more shrinkage

MSE of James-Stein Estimator

$$\begin{aligned}R(\theta, \hat{\theta}^{JS}) &= \mathbb{E} \left[(\hat{\theta}^{JS} - \theta)' (\hat{\theta}^{JS} - \theta) \right] \\&= \mathbb{E} \left[\left\{ (X - \theta) - \frac{(p-2)X}{X'X} \right\}' \left\{ (X - \theta) - \frac{(p-2)X}{X'X} \right\} \right] \\&= \mathbb{E} [(X - \theta)' (X - \theta)] - 2(p-2) \mathbb{E} \left[\frac{X'(X - \theta)}{X'X} \right] \\&\quad + (p-2)^2 \mathbb{E} \left[\frac{1}{X'X} \right] \\&= p - 2(p-2) \mathbb{E} \left[\frac{X'(X - \theta)}{X'X} \right] + (p-2)^2 \mathbb{E} \left[\frac{1}{X'X} \right]\end{aligned}$$

Using fact that $R(\theta, \hat{\theta}) = p$

Simplifying the Second Term

Writing Numerator as a Sum

$$\mathbb{E} \left[\frac{X'(X - \theta)}{X'X} \right] = \mathbb{E} \left[\frac{\sum_{i=1}^p X_i (X_i - \theta_i)}{X'X} \right] = \sum_{i=1}^p \mathbb{E} \left[\frac{X_i (X_i - \theta_i)}{X'X} \right]$$

For $i = 1, \dots, p$

$$\mathbb{E} \left[\frac{X_i (X_i - \theta_i)}{X'X} \right] = \mathbb{E} \left[\frac{X'X - 2X_i^2}{(X'X)^2} \right]$$

Not obvious: integration by parts, expectation as a p -fold integral, $X \sim N(\theta, I)$

Combining

$$\begin{aligned} \mathbb{E} \left[\frac{X'(X - \theta)}{X'X} \right] &= \sum_{i=1}^p \mathbb{E} \left[\frac{X'X - 2X_i^2}{(X'X)^2} \right] = p \mathbb{E} \left[\frac{1}{X'X} \right] - 2 \mathbb{E} \left[\frac{\sum_{i=1}^p X_i^2}{(X'X)^2} \right] \\ &= p \mathbb{E} \left[\frac{1}{X'X} \right] - 2 \mathbb{E} \left[\frac{X'X}{(X'X)^2} \right] = (p - 2) \mathbb{E} \left[\frac{1}{X'X} \right] \end{aligned}$$

The MLE is Inadmissible when $p \geq 3$

$$\begin{aligned} R\left(\theta, \hat{\theta}^{JS}\right) &= p - 2(p-2) \left\{ (p-2) \mathbb{E} \left[\frac{1}{X'X} \right] \right\} + (p-2)^2 \mathbb{E} \left[\frac{1}{X'X} \right] \\ &= p - (p-2)^2 \mathbb{E} \left[\frac{1}{X'X} \right] \end{aligned}$$

- ▶ $\mathbb{E}[1/(X'X)]$ exists and is positive whenever $p \geq 3$
- ▶ $(p-2)^2$ is always positive
- ▶ Hence, second term in the MSE expression is *negative*
- ▶ First term is MSE of the MLE

Therefore James-Stein strictly dominates MLE whenever $p \geq 3$!

James-Stein More Generally

- ▶ Our example was specific, but the result is general:
 - ▶ MLE is inadmissible under quadratic loss in regression model with at least three regressors.
 - ▶ Note, however, that this is MSE for the *full parameter vector*
- ▶ James-Stein estimator is also inadmissible!
 - ▶ Dominated by “positive-part” James-Stein estimator:

$$\hat{\beta}^{JS} = \hat{\beta} \left[1 - \frac{(p-2)\hat{\sigma}^2}{\hat{\beta}'X'X\hat{\beta}} \right]_+$$

- ▶ $\hat{\beta} = \text{OLS}$, $(x)_+ = \max(x, 0)$, $\hat{\sigma}^2 = \text{usual OLS-based estimator}$
- ▶ Stops us from shrinking *past* zero to get a negative estimate for an element of β with a small OLS estimate.
- ▶ Positive-part James-Stein isn't admissible either!

Lecture #2 – Model Selection I

Kullback-Leibler Divergence

Bias of Maximized Sample Log-Likelihood

Review of Asymptotics for Mis-specified MLE

Deriving AIC and TIC

Corrected AIC (AIC_c)

Mallow's C_p

Kullback-Leibler (KL) Divergence

Motivation

How well does a given density $f(y)$ approximate an unknown true density $g(y)$? Use this to select between parametric models.

Definition

$$\text{KL}(g; f) = \underbrace{\mathbb{E}_G \left[\log \left\{ \frac{g(Y)}{f(Y)} \right\} \right]}_{\text{True density on top}} = \underbrace{\mathbb{E}_G [\log g(Y)]}_{\substack{\text{Depends only on truth} \\ \text{Fixed across models}}} - \underbrace{\mathbb{E}_G [\log f(Y)]}_{\text{Expected log-likelihood}}$$

Properties

- ▶ Not symmetric: $\text{KL}(g; f) \neq \text{KL}(f; g)$
- ▶ By Jensen's Inequality: $\text{KL}(g; f) \geq 0$ (strict iff $g = f$ a.e.)
- ▶ Minimize KL \iff Maximize Expected log-likelihood

KL Divergence and Mis-specified MLE

Pseudo-true Parameter Value θ_0

$$\hat{\theta}_{MLE} \xrightarrow{P} \theta_0 \equiv \arg \min_{\theta \in \Theta} \text{KL}(g; f_{\theta}) = \arg \max_{\theta \in \Theta} \mathbb{E}_G[\log f(Y|\theta)]$$

What if f_{θ} is correctly specified?

If $g = f_{\theta}$ for some θ then $\text{KL}(g; f_{\theta})$ is minimized at zero.

Goal: Compare Mis-specified Models

$$\mathbb{E}_G [\log f(Y|\theta_0)] \quad \text{versus} \quad \mathbb{E}_G [\log h(Y|\gamma_0)]$$

where θ_0 is the pseudo-true parameter value for f_{θ} and γ_0 is the pseudo-true parameter value for h_{γ} .

How to Estimate Expected Log Likelihood?

For simplicity: $Y_1, \dots, Y_n \sim \text{iid } g(y)$

Unbiased but Infeasible

$$\mathbb{E}_G \left[\frac{1}{T} \ell(\theta_0) \right] = \mathbb{E}_G \left[\frac{1}{T} \sum_{t=1}^T \log f(Y_t | \theta_0) \right] = \mathbb{E}_G [\log f(Y | \theta_0)]$$

Biased but Feasible

$T^{-1} \ell(\hat{\theta}_{MLE})$ is a **biased** estimator of $\mathbb{E}_G[\log f(Y | \theta_0)]$.

Intuition for the Bias

$T^{-1} \ell(\hat{\theta}_{MLE}) > T^{-1} \ell(\theta_0)$ unless $\hat{\theta}_{MLE} = \theta_0$. Maximized sample log-like. is an **overly optimistic** estimator of expected log-like.

What to do about this bias?

1. General-purpose asymptotic approximation of “degree of over-optimism” of maximized sample log-likelihood.
 - ▶ Takeuchi's Information Criterion (TIC)
 - ▶ Akaike's Information Criterion (AIC)
2. Problem-specific finite sample approach, assuming $g \in f_\theta$.
 - ▶ Corrected AIC (AIC_c) of Hurvich and Tsai (1989)

Tradeoffs

TIC is most general and makes weakest assumptions, but requires very large T to work well. AIC is a good approximation to TIC that requires less data. Both AIC and TIC perform poorly when T is small relative to the number of parameters, hence AIC_c .

Recall: Asymptotics for Mis-specified ML Estimation

Model $f(y|\theta)$, pseudo-true parameter θ_0 . For simplicity $Y_1, \dots, Y_T \sim \text{iid } g(y)$.

Fundamental Expansion

$$\sqrt{T}(\hat{\theta} - \theta_0) = J^{-1} \left(\sqrt{T} \bar{U}_T \right) + o_p(1)$$

$$J = -\mathbb{E}_G \left[\frac{\partial \log f(Y|\theta_0)}{\partial \theta \partial \theta'} \right], \quad \bar{U}_T = \frac{1}{T} \sum_{t=1}^T \frac{\partial \log f(Y_t|\theta_0)}{\partial \theta}$$

Central Limit Theorem

$$\sqrt{T} \bar{U}_T \rightarrow_d U \sim N_p(0, K), \quad K = \text{Var}_G \left[\frac{\partial \log f(Y|\theta_0)}{\partial \theta} \right]$$

$$\sqrt{T}(\hat{\theta} - \theta_0) \rightarrow_d J^{-1} U \sim N_p(0, J^{-1} K J^{-1})$$

Information Matrix Equality

If $g = f_\theta$ for some $\theta \in \Theta$ then $K = J \implies \text{AVAR}(\hat{\theta}) = J^{-1}$

Bias Relative to Infeasible Plug-in Estimator

Definition of Bias Term B

$$B = \underbrace{\frac{1}{T} \ell(\hat{\theta})}_{\text{feasible over-optimistic}} - \underbrace{\int g(y) \log f(y|\hat{\theta}) dy}_{\text{uses data only once infeas. not over-optimistic}}$$

Question to Answer

On average, over the sampling distribution of $\hat{\theta}$, how large is B ?

AIC and TIC construct an asymptotic approximation of $\mathbb{E}[B]$.

Derivation of AIC/TIC

Step 1: Taylor Expansion

$$B = \bar{Z}_T + (\hat{\theta} - \theta_0)' J(\hat{\theta} - \theta_0) + o_p(T^{-1})$$

$$\bar{Z}_T = \frac{1}{T} \sum_{t=1}^T \{\log f(Y_t|\theta_0) - \mathbb{E}_G[\log f(Y|\theta_0)]\}$$

Step 2: $\mathbb{E}[\bar{Z}_T] = 0$

$$\mathbb{E}[B] \approx \mathbb{E} \left[(\hat{\theta} - \theta_0)' J(\hat{\theta} - \theta_0) \right]$$

Step 3: $\sqrt{T}(\hat{\theta} - \theta_0) \rightarrow_d J^{-1}U$

$$T(\hat{\theta} - \theta_0)' J(\hat{\theta} - \theta_0) \rightarrow_d U' J^{-1}U$$

Derivation of AIC/TIC Continued...

Step 3: $\sqrt{T}(\hat{\theta} - \theta_0) \rightarrow_d J^{-1}U$

$$T(\hat{\theta} - \theta_0)'J(\hat{\theta} - \theta_0) \rightarrow_d U'J^{-1}U$$

Step 4: $U \sim N_p(0, K)$

$$\mathbb{E}[B] \approx \frac{1}{T}\mathbb{E}[U'J^{-1}U] = \frac{1}{T}\text{tr}\{J^{-1}K\}$$

Final Result:

$T^{-1}\text{tr}\{J^{-1}K\}$ is an asymp. unbiased estimator of the over-optimism of $T^{-1}\ell(\hat{\theta})$ relative to $\int g(y) \log f(y|\hat{\theta}) dy$.

TIC and AIC

Takeuchi's Information Criterion

Multiply by $2T$, estimate $J, K \Rightarrow \text{TIC} = 2 \left[\ell(\hat{\theta}) - \text{tr} \left\{ \hat{J}^{-1} \hat{K} \right\} \right]$

Akaike's Information Criterion

If $g = f_{\theta}$ then $J = K \Rightarrow \text{tr} \left\{ J^{-1} K \right\} = p \Rightarrow \text{AIC} = 2 \left[\ell(\hat{\theta}) - p \right]$

Contrasting AIC and TIC

Technically, AIC requires that all models under consideration are at least correctly specified while TIC doesn't. But $J^{-1}K$ is hard to estimate, and if a model is badly mis-specified, $\ell(\hat{\theta})$ dominates.

Corrected AIC (AIC_c) – Hurvich & Tsai (1989)

Idea Behind AIC_c

Asymptotic approximation used for AIC/TIC works poorly if p is too large relative to T . Try exact, finite-sample approach instead.

Assumption: True DGP

$$\mathbf{y} = \mathbf{X}\beta_0 + \varepsilon, \quad \varepsilon \sim N(\mathbf{0}, \sigma_0^2 \mathbf{I}_T), \quad k \text{ Regressors}$$

Can Show That

$$KL(g, f) = \frac{T}{2} \left[\frac{\sigma_0^2}{\sigma_1^2} - \log \left(\frac{\sigma_0^2}{\sigma_1^2} \right) - 1 \right] + \left(\frac{1}{2\sigma_1^2} \right) (\beta_0 - \beta_1)' \mathbf{X}' \mathbf{X} (\beta_0 - \beta_1)$$

Where f is a normal regression model with parameters (β_1, σ_1^2) that might not be the true parameters.

But how can we use this?

$$KL(g, f) = \frac{T}{2} \left[\frac{\sigma_0^2}{\sigma_1^2} - \log \left(\frac{\sigma_0^2}{\sigma_1^2} \right) - 1 \right] + \left(\frac{1}{2\sigma_1^2} \right) (\beta_0 - \beta_1)' \mathbf{X}' \mathbf{X} (\beta_0 - \beta_1)$$

1. Would need to know (β_1, σ_1^2) for **candidate model**.
 - ▶ Easy: just use MLE $(\hat{\beta}_1, \hat{\sigma}_1^2)$
2. Would need to know (β_0, σ_0^2) for **true model**.
 - ▶ Very hard! The whole problem is that we don't know these!

Hurvich & Tsai (1989) Assume:

- ▶ Every candidate model is **at least correctly specified**
- ▶ Implies any candidate estimator $(\hat{\beta}, \hat{\sigma}^2)$ is consistent for truth.

Deriving the Corrected AIC

Since $(\hat{\beta}, \hat{\sigma}^2)$ are random, look at $\mathbb{E}[\widehat{KL}]$, where

$$\widehat{KL} = \frac{T}{2} \left[\frac{\sigma_0^2}{\hat{\sigma}^2} - \log \left(\frac{\sigma_0^2}{\hat{\sigma}^2} \right) - 1 \right] + \left(\frac{1}{2\hat{\sigma}^2} \right) (\hat{\beta} - \beta_0)' \mathbf{X}' \mathbf{X} (\hat{\beta} - \beta_0)$$

Finite-sample theory for correctly spec. normal regression model:

$$\mathbb{E}[\widehat{KL}] = \frac{T}{2} \left\{ \frac{T+k}{T-k-2} - \log(\sigma_0^2) + \mathbb{E}[\log \hat{\sigma}^2] - 1 \right\}$$

Eliminate constants and scaling, unbiased estimator of $\mathbb{E}[\log \hat{\sigma}^2]$:

$$\text{AIC}_c = \log \hat{\sigma}^2 + \frac{T+k}{T-k-2}$$

a finite-sample unbiased estimator of KL for model comparison

Motivation: Predict \mathbf{y} from \mathbf{x} via Linear Regression

$$\underset{(T \times 1)}{\mathbf{y}} = \underset{(T \times K)}{\mathbf{X}} \underset{(K \times 1)}{\boldsymbol{\beta}} + \boldsymbol{\epsilon}$$

$$\mathbb{E}[\boldsymbol{\epsilon}|\mathbf{X}] = 0, \quad \text{Var}(\boldsymbol{\epsilon}|\mathbf{X}) = \sigma^2 \mathbf{I}$$

- ▶ If $\boldsymbol{\beta}$ were known, could never achieve lower MSE than by using all regressors to predict.
- ▶ But $\boldsymbol{\beta}$ is unknown so we have to estimate it from data \Rightarrow bias-variance tradeoff.
- ▶ Could make sense to exclude regressors with small coefficients: add small bias but reduce variance.

Operationalizing the Bias-Variance Tradeoff Idea

Mallow's C_p

Approximate the predictive MSE of each model relative to the infeasible optimum in which β is known.

Notation

- ▶ Model index m and regressor matrix \mathbf{X}_m
- ▶ Corresponding OLS estimator $\hat{\beta}_m$ padded out with zeros
- ▶ $\mathbf{X}\hat{\beta}_m = \mathbf{X}_{(-m)}\mathbf{0} + \mathbf{X}_m [(\mathbf{X}_m'\mathbf{X}_m)^{-1}\mathbf{X}_m'\mathbf{y}] = \mathbf{P}_m\mathbf{y}$

In-sample versus Out-of-sample Prediction Error

Why not compare $RSS(m)$?

In-sample prediction error: $RSS(m) = (\mathbf{y} - \mathbf{X}\hat{\beta}_m)'(\mathbf{y} - \mathbf{X}\hat{\beta}_m)$

From your Problem Set

RSS cannot decrease even if we add irrelevant regressors. Thus in-sample prediction error is an **overly optimistic** estimate of out-of-sample prediction error.

Bias-Variance Tradeoff

Out-of-sample performance of full model (using all regressors) could be very poor if there is a lot of estimation uncertainty associated with regressors that aren't very predictive.

Predictive MSE of $\mathbf{X}\hat{\boldsymbol{\beta}}_m$ relative to infeasible optimum $\mathbf{X}\boldsymbol{\beta}$

Step 1: Algebra

$$\begin{aligned}\mathbf{X}\hat{\boldsymbol{\beta}}_m - \mathbf{X}\boldsymbol{\beta} &= \mathbf{P}_m\mathbf{y} - \mathbf{X}\boldsymbol{\beta} = \mathbf{P}_m(\mathbf{y} - \mathbf{X}\boldsymbol{\beta}) - (\mathbf{I} - \mathbf{P}_m)\mathbf{X}\boldsymbol{\beta} \\ &= \mathbf{P}_m\boldsymbol{\epsilon} - (\mathbf{I} - \mathbf{P}_m)\mathbf{X}\boldsymbol{\beta}\end{aligned}$$

Step 2: \mathbf{P}_m and $(\mathbf{I} - \mathbf{P}_m)$ are both symmetric and idempotent, and orthogonal to each other

$$\begin{aligned}\left\|\mathbf{X}\hat{\boldsymbol{\beta}}_m - \mathbf{X}\boldsymbol{\beta}\right\|^2 &= \{\mathbf{P}_m\boldsymbol{\epsilon} - (\mathbf{I} - \mathbf{P}_m)\mathbf{X}\boldsymbol{\beta}\}' \{\mathbf{P}_m\boldsymbol{\epsilon} + (\mathbf{I} - \mathbf{P}_m)\mathbf{X}\boldsymbol{\beta}\} \\ &= \boldsymbol{\epsilon}'\mathbf{P}_m'\mathbf{P}_m\boldsymbol{\epsilon} - \boldsymbol{\beta}'\mathbf{X}'(\mathbf{I} - \mathbf{P}_m)'\mathbf{P}_m\boldsymbol{\epsilon} - \boldsymbol{\epsilon}'\mathbf{P}_m'(\mathbf{I} - \mathbf{P}_m)\mathbf{X}\boldsymbol{\beta} \\ &\quad + \boldsymbol{\beta}'\mathbf{X}'(\mathbf{I} - \mathbf{P}_m)(\mathbf{I} - \mathbf{P}_m)\mathbf{X}\boldsymbol{\beta} \\ &= \boldsymbol{\epsilon}'\mathbf{P}_m\boldsymbol{\epsilon} + \boldsymbol{\beta}'\mathbf{X}'(\mathbf{I} - \mathbf{P}_m)\mathbf{X}\boldsymbol{\beta}\end{aligned}$$

Predictive MSE of $\mathbf{X}\hat{\boldsymbol{\beta}}_m$ relative to infeasible optimum $\mathbf{X}\boldsymbol{\beta}$

Step 3: Expectation of Step 2 conditional on \mathbf{X}

$$\begin{aligned}\text{MSE}(m|\mathbf{X}) &= \mathbb{E} \left[(\mathbf{X}\hat{\boldsymbol{\beta}}_m - \mathbf{X}\boldsymbol{\beta})' (\mathbf{X}\hat{\boldsymbol{\beta}}_m - \mathbf{X}\boldsymbol{\beta}) | \mathbf{X} \right] \\&= \mathbb{E} \left[\boldsymbol{\epsilon}' \mathbf{P}_m \boldsymbol{\epsilon} | \mathbf{X} \right] + \mathbb{E} \left[\boldsymbol{\beta}' \mathbf{X}' (\mathbf{I} - \mathbf{P}_m) \mathbf{X} \boldsymbol{\beta} | \mathbf{X} \right] \\&= \mathbb{E} \left[\text{tr} \left\{ \boldsymbol{\epsilon}' \mathbf{P}_m \boldsymbol{\epsilon} \right\} | \mathbf{X} \right] + \boldsymbol{\beta}' \mathbf{X}' (\mathbf{I} - \mathbf{P}_m) \mathbf{X} \boldsymbol{\beta} \\&= \text{tr} \left\{ \mathbb{E} [\boldsymbol{\epsilon} \boldsymbol{\epsilon}' | \mathbf{X}] \mathbf{P}_m \right\} + \boldsymbol{\beta}' \mathbf{X}' (\mathbf{I} - \mathbf{P}_m) \mathbf{X} \boldsymbol{\beta} \\&= \text{tr} \left\{ \sigma^2 \mathbf{P}_m \right\} + \boldsymbol{\beta}' \mathbf{X}' (\mathbf{I} - \mathbf{P}_m) \mathbf{X} \boldsymbol{\beta} \\&= \sigma^2 k_m + \boldsymbol{\beta}' \mathbf{X}' (\mathbf{I} - \mathbf{P}_m) \mathbf{X} \boldsymbol{\beta}\end{aligned}$$

where k_m denotes the number of regressors in \mathbf{X}_m and

$$\text{tr}(\mathbf{P}_m) = \text{tr} \left\{ \mathbf{X}_m (\mathbf{X}_m' \mathbf{X}_m)^{-1} \mathbf{X}_m' \right\} = \text{tr} \left\{ \mathbf{X}_m' \mathbf{X}_m (\mathbf{X}_m' \mathbf{X}_m)^{-1} \right\} = \text{tr}(\mathbf{I}_m)$$

Now we know the MSE of a given model...

$$\text{MSE}(m|\mathbf{X}) = \sigma^2 k_m + \beta' \mathbf{X}' (\mathbf{I} - \mathbf{P}_m) \mathbf{X} \beta$$

Bias-Variance Tradeoff

- ▶ Smaller Model $\Rightarrow \sigma^2 k_m$ smaller: less estimation uncertainty.
- ▶ Bigger Model $\Rightarrow \mathbf{X}' (\mathbf{I} - \mathbf{P}_m) \mathbf{X} = \|(\mathbf{I} - \mathbf{P}_m) \mathbf{X}\|^2$ is in general smaller: less (squared) bias.

Mallow's C_p

- ▶ Problem: MSE formula is infeasible since it involves β and σ^2 .
- ▶ Solution: Mallow's C_p constructs an unbiased estimator.
- ▶ Idea: what about plugging in $\hat{\beta}$ to estimate second term?

What if we plug in $\hat{\beta}$ to estimate the second term?

For the missing algebra in Step 4, see the lecture notes.

Notation

Let $\hat{\beta}$ denote the full model estimator and \mathbf{P} be the corresponding projection matrix: $\mathbf{X}\hat{\beta} = \mathbf{P}\mathbf{y}$.

Crucial Fact

$\text{span}(\mathbf{X}_m)$ is a subspace of $\text{span}(\mathbf{X})$, so $\mathbf{P}_m\mathbf{P} = \mathbf{P}\mathbf{P}_m = \mathbf{P}_m$.

Step 4: Algebra using the preceding fact

$$\mathbb{E} \left[\hat{\beta}' \mathbf{X}' (\mathbf{I} - \mathbf{P}_m) \mathbf{X} \hat{\beta} | \mathbf{X} \right] = \dots = \beta' \mathbf{X}' (\mathbf{I} - \mathbf{P}_m) \mathbf{X} \beta + \mathbb{E} \left[\epsilon' (\mathbf{P} - \mathbf{P}_m) \epsilon | \mathbf{X} \right]$$

Substituting $\hat{\beta}$ doesn't work...

Step 5: Use “Trace Trick” on second term from Step 4

$$\begin{aligned}\mathbb{E}[\epsilon'(\mathbf{P} - \mathbf{P}_m)\epsilon|\mathbf{X}] &= \mathbb{E}[\text{tr}\{\epsilon'(\mathbf{P} - \mathbf{P}_m)\epsilon\}|\mathbf{X}] \\ &= \text{tr}\{\mathbb{E}[\epsilon\epsilon'|\mathbf{X}](\mathbf{P} - \mathbf{P}_m)\} \\ &= \text{tr}\{\sigma^2(\mathbf{P} - \mathbf{P}_m)\} \\ &= \sigma^2(\text{trace}\{\mathbf{P}\} - \text{trace}\{\mathbf{P}_m\}) \\ &= \sigma^2(K - k_m)\end{aligned}$$

where K is the total number of regressors in \mathbf{X}

Bias of Plug-in Estimator

$$\mathbb{E}\left[\hat{\beta}'\mathbf{X}'(\mathbf{I} - \mathbf{P}_m)\mathbf{X}\hat{\beta}|\mathbf{X}\right] = \underbrace{\beta'\mathbf{X}'(\mathbf{I} - \mathbf{P}_m)\mathbf{X}\beta}_{\text{Truth}} + \underbrace{\sigma^2(K - k_m)}_{\text{Bias}}$$

Putting Everything Together: Mallows's C_p

Want An Unbiased Estimator of This:

$$\text{MSE}(m|\mathbf{X}) = \sigma^2 k_m + \beta' \mathbf{X}' (\mathbf{I} - \mathbf{P}_m) \mathbf{X} \beta$$

Previous Slide:

$$\mathbb{E} \left[\hat{\beta}' \mathbf{X}' (\mathbf{I} - \mathbf{P}_m) \mathbf{X} \hat{\beta} | \mathbf{X} \right] = \beta' \mathbf{X}' (\mathbf{I} - \mathbf{P}_m) \mathbf{X} \beta + \sigma^2 (K - k_m)$$

End Result:

$$\begin{aligned} \text{MC}(m) &= \hat{\sigma}^2 k_m + \left[\hat{\beta}' \mathbf{X}' (\mathbf{I} - \mathbf{P}_m) \mathbf{X} \hat{\beta} - \hat{\sigma}^2 (K - k_m) \right] \\ &= \hat{\beta}' \mathbf{X}' (\mathbf{I} - \mathbf{P}_m) \mathbf{X} \hat{\beta} + \hat{\sigma}^2 (2k_m - K) \end{aligned}$$

is an unbiased estimator of MSE, with $\hat{\sigma}^2 = \mathbf{y}'(\mathbf{I} - \mathbf{P})\mathbf{y}/(T - K)$

Why is this different from the textbook formula?

Just algebra, but tedious. . .

$$\begin{aligned}\text{MC}(m) - 2\hat{\sigma}^2 k_m &= \hat{\beta}' X' (\mathbf{I} - P_M) X \hat{\beta} - K \hat{\sigma}^2 \\ &\vdots \\ &= \mathbf{y}' (\mathbf{I} - P_M) \mathbf{y} - T \hat{\sigma}^2 \\ &= \text{RSS}(m) - T \hat{\sigma}^2\end{aligned}$$

Therefore:

$$\text{MC}(m) = \text{RSS}(m) + \hat{\sigma}^2(2k_m - T)$$

Divide Through by $\hat{\sigma}^2$:

$$C_p(m) = \frac{\text{RSS}(m)}{\hat{\sigma}^2} + 2k_m - T$$

Tells us how to adjust RSS for number of regressors. . .

Lecture #3 – Model Selection II

Bayesian Model Comparison

Bayesian Information Criterion (BIC)

K-fold Cross-validation

Asymptotic Equivalence Between LOO-CV and TIC

Bayesian Model Comparison: Marginal Likelihoods

Bayes' Rule for Model $m \in \mathcal{M}$

$$\underbrace{\pi(\boldsymbol{\theta}|\mathbf{y}, m)}_{\text{Posterior}} \propto \underbrace{\pi(\boldsymbol{\theta}|m)}_{\text{Prior}} \underbrace{f(\mathbf{y}|\boldsymbol{\theta}, m)}_{\text{Likelihood}}$$
$$\underbrace{f(\mathbf{y}|m)}_{\text{Marginal Likelihood}} = \int_{\Theta} \pi(\boldsymbol{\theta}|m) f(\mathbf{y}|\boldsymbol{\theta}, m) \, d\boldsymbol{\theta}$$

Posterior Model Probability for $m \in \mathcal{M}$

$$P(m|\mathbf{y}) = \frac{P(m)f(\mathbf{y}|m)}{f(\mathbf{y})} = \frac{\int_{\Theta} P(m)f(\mathbf{y}, \boldsymbol{\theta}|m) \, d\boldsymbol{\theta}}{f(\mathbf{y})} = \frac{P(m)}{f(\mathbf{y})} \int_{\Theta} \pi(\boldsymbol{\theta}|m)f(\mathbf{y}|\boldsymbol{\theta}, m) \, d\boldsymbol{\theta}$$

where $P(m)$ is the **prior model probability** and $f(\mathbf{y})$ is constant across models.

Laplace (aka Saddlepoint) Approximation

Suppress model index m for simplicity.

General Case: for T large...

$$\int_{\Theta} g(\boldsymbol{\theta}) \exp\{T \cdot h(\boldsymbol{\theta})\} d\boldsymbol{\theta} \approx \left(\frac{2\pi}{T}\right)^{p/2} \exp\{T \cdot h(\boldsymbol{\theta}_0)\} g(\boldsymbol{\theta}_0) |H(\boldsymbol{\theta}_0)|^{-1/2}$$

$$p = \dim(\boldsymbol{\theta}), \quad \boldsymbol{\theta}_0 = \arg \max_{\boldsymbol{\theta} \in \Theta} h(\boldsymbol{\theta}), \quad H(\boldsymbol{\theta}_0) = -\frac{\partial^2 h(\boldsymbol{\theta})}{\partial \boldsymbol{\theta} \partial \boldsymbol{\theta}'} \Big|_{\boldsymbol{\theta}=\boldsymbol{\theta}_0}$$

Use to Approximate Marginal Likelihood

$$h(\boldsymbol{\theta}) = \frac{\ell(\boldsymbol{\theta})}{T} = \frac{1}{T} \sum_{t=1}^T \log f(Y_t | \boldsymbol{\theta}), \quad H(\boldsymbol{\theta}) = J_T(\boldsymbol{\theta}) = -\frac{1}{T} \sum_{t=1}^T \frac{\partial^2 \log f(Y_t | \boldsymbol{\theta})}{\partial \boldsymbol{\theta} \partial \boldsymbol{\theta}'}, \quad g(\boldsymbol{\theta}) = \pi(\boldsymbol{\theta})$$

and substitute $\hat{\boldsymbol{\theta}}_{MLE}$ for $\boldsymbol{\theta}_0$

Laplace Approximation to Marginal Likelihood

Suppress model index m for simplicity.

$$\int_{\Theta} \pi(\boldsymbol{\theta}) f(\mathbf{y}|\boldsymbol{\theta}) \, d\boldsymbol{\theta} \approx \left(\frac{2\pi}{T}\right)^{p/2} \exp\left\{\ell(\hat{\boldsymbol{\theta}}_{MLE})\right\} \pi(\hat{\boldsymbol{\theta}}_{MLE}) \left|J_T(\hat{\boldsymbol{\theta}}_{MLE})\right|^{-1/2}$$

$$\ell(\boldsymbol{\theta}) = \sum_{t=1}^T \log f(Y_t|\boldsymbol{\theta}), \quad H(\boldsymbol{\theta}) = J_T(\boldsymbol{\theta}) = -\frac{1}{T} \sum_{t=1}^T \frac{\partial^2 \log f(Y_t|\boldsymbol{\theta})}{\partial \boldsymbol{\theta} \partial \boldsymbol{\theta}'}$$

Bayesian Information Criterion

$$\int_{\Theta} \pi(\boldsymbol{\theta}) f(\mathbf{y}|\boldsymbol{\theta}) \, d\boldsymbol{\theta} \approx \left(\frac{2\pi}{T}\right)^{p/2} \exp\left\{\ell(\hat{\boldsymbol{\theta}}_{MLE})\right\} \pi(\hat{\boldsymbol{\theta}}_{MLE}) \left|J_T(\hat{\boldsymbol{\theta}}_{MLE})\right|^{-1/2}$$

Take Logs and Multiply by 2

$$2 \log f(\mathbf{y}|\boldsymbol{\theta}) \approx \underbrace{2\ell(\hat{\boldsymbol{\theta}}_{MLE})}_{O_p(T)} - \underbrace{p \log(T)}_{O(\log T)} + \underbrace{p \log(2\pi) + \log \pi(\hat{\boldsymbol{\theta}}) - \log |J_T(\hat{\boldsymbol{\theta}})|}_{O_p(1)}$$

The BIC

Assume uniform prior over **models** and ignore lower order terms:

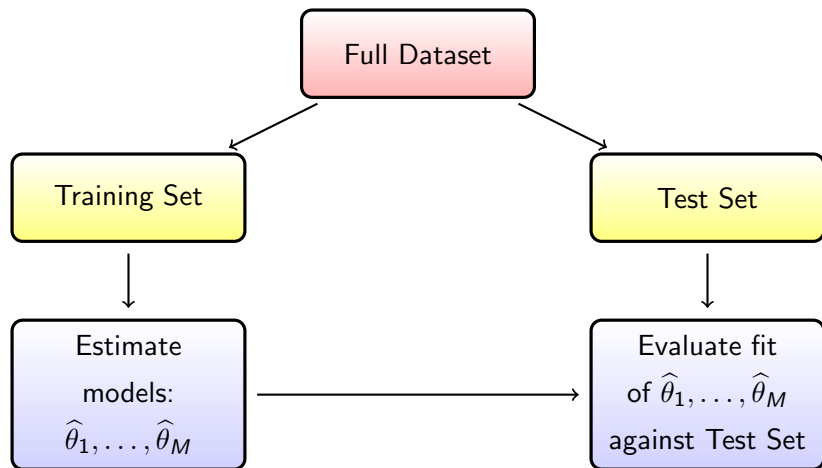
$$\text{BIC}(m) = 2 \log f(\mathbf{y}|\hat{\boldsymbol{\theta}}, m) - p_m \log(T)$$

large-sample Frequentist approx. to Bayesian marginal likelihood

Model Selection using a Hold-out Sample

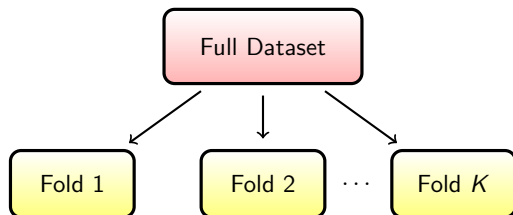
- ▶ The real problem is **double** use of the data: first for estimation, then for model comparison.
 - ▶ Maximized sample log-likelihood is an overly optimistic estimate of expected log-likelihood and hence KL-divergence
 - ▶ In-sample squared prediction error is an overly optimistic estimator of out-of-sample squared prediction error
- ▶ AIC/TIC, AIC_c , BIC, C_p **penalize** sample log-likelihood or RSS to compensate.
- ▶ Another idea: **don't re-use the same data!**

Hold-out Sample: Partition the Full Dataset



Unfortunately this is extremely wasteful of data...

K-fold Cross-Validation: “Pseudo-out-of-sample”



Step 1

Randomly partition full dataset into K folds of approx. equal size.

Step 2

Treat k^{th} fold as a hold-out sample and estimate model using all observations **except** those in fold k : yielding estimator $\hat{\theta}(-k)$.

K -fold Cross-Validation: “Pseudo-out-of-sample”

Step 2

Treat k^{th} fold as a hold-out sample and estimate model using all observations **except** those in fold k : yielding estimator $\hat{\theta}(-k)$.

Step 3

Repeat Step 2 for each $k = 1, \dots, K$.

Step 4

For each t calculate the prediction $\hat{y}_t^{-k(t)}$ of y_t based on $\hat{\theta}(-k(t))$, the estimator that excluded observation t .

K -fold Cross-Validation: “Pseudo-out-of-sample”

Step 4

For each t calculate the prediction $\hat{y}_t^{-k(t)}$ of y_t based on $\hat{\theta}(-k(t))$, the estimator that excluded observation t .

Step 5

Define $CV_K = \frac{1}{T} \sum_{t=1}^T L(y_t, \hat{y}_t^{-k(t)})$ where L is a loss function.

Step 5

Repeat for each model & choose m to minimize $CV_K(m)$.

CV uses each observation for parameter estimation and model evaluation but never at the same time!

Cross-Validation (CV): Some Details

Which Loss Function?

- ▶ For regression squared error loss makes sense
- ▶ For classification (discrete prediction) could use zero-one loss.
- ▶ Can also use log-likelihood/KL-divergence as a loss function. . .

How Many Folds?

- ▶ One extreme: $K = 2$. Closest to Training/Test idea.
- ▶ Other extreme: $K = T$ **Leave-one-out** CV (LOO-CV).
- ▶ Computationally expensive model \Rightarrow may prefer fewer folds.
- ▶ If your model is a linear smoother there's a computational trick that makes LOO-CV extremely fast. (Problem Set)
- ▶ Asymptotic properties are related to K . . .

Relationship between LOO-CV and TIC

Theorem

LOO-CV using KL-divergence as the loss function is asymptotically equivalent to TIC but doesn't require us to estimate the Hessian and variance of the score.

Large-sample Equivalence of LOO-CV and TIC

Notation and Assumptions

For simplicity let $Y_1, \dots, Y_T \sim \text{iid}$. Let $\hat{\theta}_{(t)}$ be the maximum likelihood estimator based on all observations **except** t and $\hat{\theta}$ be the full-sample estimator.

Log-likelihood as “Loss”

$CV_1 = \frac{1}{T} \sum_{t=1}^T \log f(y_t | \hat{\theta}_{(t)})$ but since min. KL = max. log-like.
we choose the model with **highest** $CV_1(m)$.

Overview of the Proof

First-Order Taylor Expansion of $\log f(y_t|\hat{\theta}_{(t)})$ around $\hat{\theta}$:

$$\begin{aligned} CV_1 &= \frac{1}{T} \sum_{t=1}^T \log f(y_t|\hat{\theta}_{(t)}) \\ &= \frac{1}{T} \sum_{t=1}^T \left[\log f(y_t|\hat{\theta}) + \frac{\partial \log f(y_t|\hat{\theta})}{\partial \theta'} (\hat{\theta}_{(t)} - \hat{\theta}) \right] + o_p(1) \\ &= \frac{\ell(\hat{\theta})}{T} + \frac{1}{T} \sum_{t=1}^T \frac{\partial \log f(y_t|\hat{\theta})}{\partial \theta'} (\hat{\theta}_{(t)} - \hat{\theta}) + o_p(1) \end{aligned}$$

Why isn't the first-order term zero in this case?

Important Side Point

Definition of ML Estimator

$$\frac{\partial \ell(\hat{\theta})}{\partial \theta'} = \frac{1}{T} \sum_{t=1}^T \frac{\partial \log f(y_t | \hat{\theta})}{\partial \theta} = 0$$

In Contrast

$$\begin{aligned} \frac{1}{T} \sum_{t=1}^T \frac{\partial \log f(y_t | \hat{\theta})}{\partial \theta'} (\hat{\theta}_{(t)} - \hat{\theta}) &= \left[\frac{1}{T} \sum_{t=1}^T \frac{\partial \log f(y_t | \hat{\theta})}{\partial \theta'} \hat{\theta}_{(t)} \right] - \hat{\theta} \left[\frac{1}{T} \sum_{t=1}^T \frac{\partial \log f(y_t | \hat{\theta})}{\partial \theta'} \right] \\ &= \frac{1}{T} \sum_{t=1}^T \frac{\partial \log f(y_t | \hat{\theta})}{\partial \theta'} \hat{\theta}_{(t)} \neq 0 \end{aligned}$$

Overview of Proof

From expansion two slides back, we simply need to show that:

$$\frac{1}{T} \sum_{t=1}^T \frac{\partial \log f(y_t | \hat{\theta})}{\partial \theta'} \left(\hat{\theta}_{(t)} - \hat{\theta} \right) = -\frac{1}{T} \text{tr} \left(\hat{J}^{-1} \hat{K} \right) + o_p(1)$$

$$\hat{K} = \frac{1}{T} \sum_{t=1}^T \left(\frac{\partial \log f(y_t | \hat{\theta})}{\partial \theta} \right) \left(\frac{\partial \log f(y_t | \hat{\theta})}{\partial \theta} \right)'$$

$$\hat{J} = -\frac{1}{T} \sum_{t=1}^T \frac{\partial^2 \log f(y_t | \hat{\theta})}{\partial \theta \partial \theta'}$$

Overview of Proof

By the definition of \hat{K} and the properties of the trace operator:

$$\begin{aligned}-\frac{1}{T}\text{tr}\left\{\hat{J}^{-1}\hat{K}\right\} &= -\frac{1}{T}\text{tr}\left\{\hat{J}^{-1}\left[\frac{1}{T}\sum_{t=1}^T\left(\frac{\partial \log f(y_t|\hat{\theta})}{\partial \theta}\right)\left(\frac{\partial \log f(y_t|\hat{\theta})}{\partial \theta}\right)'\right]\right\} \\&= \left[\frac{1}{T}\sum_{t=1}^T\text{tr}\left\{\frac{-\hat{J}^{-1}}{T}\left(\frac{\partial \log f(y_t|\hat{\theta})}{\partial \theta}\right)\left(\frac{\partial \log f(y_t|\hat{\theta})}{\partial \theta}\right)'\right\}\right] \\&= \frac{1}{T}\sum_{t=1}^T\frac{\partial \log f(y_t|\hat{\theta})}{\partial \theta'}\left(-\frac{1}{T}\hat{J}^{-1}\right)\frac{\partial \log f(y_t|\hat{\theta})}{\partial \theta}\end{aligned}$$

So it suffices to show that

$$\left(\hat{\theta}_{(t)} - \hat{\theta}\right) = -\frac{1}{T}\hat{J}^{-1}\left[\frac{\partial \log f(y_t|\hat{\theta})}{\partial \theta}\right] + o_p(1)$$

What is an Influence Function?

Statistical Functional

$\mathbb{T} = \mathbb{T}(G)$ maps a CDF G to \mathbb{R}^p .

Example: ML Estimation

$$\theta_0 = \mathbb{T}(G) = \arg \min_{\theta \in \Theta} E_G \left[\log \left\{ \frac{g(Y)}{f(Y|\theta)} \right\} \right]$$

Influence Function

Let δ_y be a **point mass** at y : $\delta_y(y) = 1$, $\delta_y(y') = 0$ for $y' \neq y$.

Influence function = functional derivative: how does a small change in G affect \mathbb{T} ?

$$\text{infl}(G, y) = \lim_{\epsilon \rightarrow 0} \frac{\mathbb{T}[(1 - \epsilon) G + \epsilon \delta_y] - \mathbb{T}(G)}{\epsilon}$$

Relating Influence Functions to $\hat{\theta}_{(t)}$

Empirical CDF \hat{G}

$$\hat{G}(a) = \frac{1}{T} \sum_{t=1}^T \mathbf{1}\{y_t \leq a\} = \frac{1}{T} \sum_{t=1}^T \delta_{y_t}(a)$$

Relation to “LOO” Empirical CDF $\hat{G}_{(t)}$

$$\hat{G} = \left(1 - \frac{1}{T}\right) \hat{G}_{(t)} + \frac{\delta_{y_t}}{T}$$

Applying \mathbb{T} to both sides...

$$\mathbb{T}(\hat{G}) = \mathbb{T}\left((1 - 1/T)\hat{G}_{(t)} + \delta_{y_t}/T\right)$$

Relating Influence Functions to $\hat{\theta}_{(t)}$

Some algebra, followed by taking $\varepsilon = 1/T$ to zero gives:

$$\mathbb{T}(\hat{G}) = \mathbb{T}\left((1 - 1/T)\hat{G}_{(t)} + \delta_{y_t}/T\right)$$

$$\mathbb{T}(\hat{G}) - \mathbb{T}(\hat{G}_{(t)}) = \mathbb{T}\left((1 - 1/T)\hat{G}_{(t)} + \delta_{y_t}/T\right) - \mathbb{T}(\hat{G}_{(t)})$$

$$\mathbb{T}(\hat{G}) - \mathbb{T}(\hat{G}_{(t)}) = \frac{1}{T} \left[\frac{\mathbb{T}\left((1 - 1/T)\hat{G}_{(t)} + \delta_{y_t}/T\right) - \mathbb{T}(\hat{G}_{(t)})}{1/T} \right]$$

$$\mathbb{T}(\hat{G}) - \mathbb{T}(\hat{G}_{(t)}) = \frac{1}{T} \text{infl}\left(\hat{G}_{(t)}, y_t\right) + o_p(1)$$

$$\hat{\theta} - \hat{\theta}_{(t)} = \frac{1}{T} \text{infl}\left(\hat{G}, y_t\right) + o_p(1)$$

Last step: difference between having \hat{G} vs. $\hat{G}_{(t)}$ in infl is negligible

Steps for Last part of TIC/LOO-CV Equivalence Proof

Step 1

Let \hat{G} denote the empirical CDF based on y_1, \dots, y_T . Then:

$$\left(\hat{\theta}_{(t)} - \hat{\theta}\right) = -\frac{1}{T} \text{infl}(\hat{G}, y_t) + o_p(1)$$

Step 2

For ML estimation it turns out that:

$$\text{infl}(G, y) = J^{-1} \frac{\partial}{\partial \theta} \log f(y|\theta_0).$$

Step 3

Evaluating Step 2 at \hat{G} and substituting into Step 2

$$\left(\hat{\theta}_{(t)} - \hat{\theta}\right) = -\frac{1}{T} \hat{J}^{-1} \left[\frac{\partial \log f(y_t|\hat{\theta})}{\partial \theta} \right] + o_p(1)$$

Derivation of Influence Function for MLE

I'll do this on the blackboard if we have time. . .

Lecture #4 – Asymptotic Properties

Overview

Weak Consistency

Consistency

Efficiency

AIC versus BIC in a Simple Example

Overview

Asymptotic Properties

What happens as the sample size increases?

Consistency

Choose “best” model with probability approaching 1 in the limit.

Efficiency

Post-model selection estimator with low risk.

Some References

Sin and White (1992, 1996), Pötscher (1991), Leeb & Pötscher (2005), Yang (2005) and Yang (2007).

Penalizing the Likelihood

Examples we've seen:

$$TIC = 2\ell_T(\hat{\theta}) - \text{trace} \left\{ \hat{J}^{-1} \hat{K} \right\}$$

$$AIC = 2\ell_T(\hat{\theta}) - 2 \text{ length}(\theta)$$

$$BIC = 2\ell_T(\hat{\theta}) - \log(T) \text{ length}(\theta)$$

Generic penalty $c_{T,k}$

$$IC(M_k) = 2 \sum_{t=1}^T \log f_{k,t}(Y_t | \hat{\theta}_k) - c_{T,k}$$

How does choice of $c_{T,k}$ affect behavior of the criterion?

Weak Consistency: Suppose M_{k_0} Uniquely Minimizes KL

Assumption

$$\liminf_{T \rightarrow \infty} \left(\min_{k \neq k_0} \frac{1}{T} \sum_{t=1}^T \{KL(g; f_{k,t}) - KL(g; f_{k_0,t})\} \right) > 0$$

Consequences

- ▶ Any criterion with $c_{T,k} > 0$ and $c_{T,k} = o_p(T)$ is weakly consistent: **selects M_{k_0} wpa 1 in the limit.**
- ▶ Weak consistency still holds if $c_{T,k}$ is zero for one of the models, so long as it is strictly positive for all the others.

Both AIC and BIC are Weakly Consistent

Both satisfy $T^{-1}c_{T,k} \xrightarrow{P} 0$.

BIC Penalty: $c_{T,k} = \log(T) \times \text{length}(\theta_k)$

AIC Penalty: $c_{T,k} = 2 \times \text{length}(\theta_k)$

Consistency: No Unique KL-minimizer

Example

If the truth is an AR(5) model then AR(6), AR(7), AR(8), etc. models **all have zero KL-divergence**.

Principle of Parsimony

Among the KL-minimizers, choose the **simplest model**, i.e. the one with the fewest parameters.

Notation

\mathcal{J} = be the set of all models that attain minimum KL-divergence

\mathcal{J}_0 = subset with the minimum number of parameters.

Sufficient Conditions for Consistency

Consistency: Select Model from \mathcal{J}_0 wpa 1

$$\lim_{T \rightarrow \infty} \mathbb{P} \left\{ \min_{\ell \in \mathcal{J} \setminus \mathcal{J}_0} [IC(M_{j_0}) - IC(M_\ell)] > 0 \right\} = 1$$

Sufficient Conditions

(i) For all $k \neq \ell \in \mathcal{J}$

$$\sum_{t=1}^T [\log f_{k,t}(Y_t | \theta_k^*) - \log f_{\ell,t}(Y_t | \theta_\ell^*)] = O_p(1)$$

where θ_k^* and θ_ℓ^* are the KL minimizing parameter values.

(ii) For all $j_0 \in \mathcal{J}_0$ and $\ell \in (\mathcal{J} \setminus \mathcal{J}_0)$

$$P(c_{T,\ell} - c_{T,j_0} \rightarrow \infty) = 1$$

BIC is Consistent; AIC and TIC Are Not

- ▶ AIC and TIC *cannot* satisfy (ii) since $(c_{T,\ell} - c_{T,j_0})$ *does not depend on sample size*.
- ▶ It turns out that AIC and TIC are *not* consistent.
- ▶ BIC is consistent:

$$c_{T,\ell} - c_{T,j_0} = \log(T) \{ \text{length}(\theta_\ell) - \text{length}(\theta_{j_0}) \}$$

- ▶ Term in braces is *positive* since $\ell \in \mathcal{J} \setminus \mathcal{J}_0$, i.e. ℓ is not as parsimonious as j_0
- ▶ $\log(T) \rightarrow \infty$, so BIC always selects a model in \mathcal{J}_0 in the limit.

Efficiency: Risk Properties of Post-selection Estimator

Setup

- ▶ Models M_0 and M_1 ; corresponding estimators $\hat{\theta}_{0,T}$ and $\hat{\theta}_{1,T}$
- ▶ Model Selection: If $\hat{M} = 0$ choose M_0 ; if $\hat{M} = 1$ choose M_1 .

Post-selection Estimator

$$\hat{\theta}_{\hat{M},T} \equiv \mathbf{1}_{\{\hat{M}=0\}} \hat{\theta}_{0,T} + \mathbf{1}_{\{\hat{M}=1\}} \hat{\theta}_{1,T}$$

Two Sources of Randomness

Variability in $\hat{\theta}_{\hat{M},T}$ arises both from $(\hat{\theta}_{0,T}, \hat{\theta}_{1,T})$ and from \hat{M} .

Question

How does the risk of $\hat{\theta}_{\hat{M},T}$ compare to that of other estimators?

Efficiency: Risk Properties of Post-selection Estimator

Pointwise-risk Adaptivity

$\hat{\theta}_{\hat{M},T}$ is **pointwise-risk adaptive** if for any fixed $\theta \in \Theta$,

$$\frac{R(\theta, \hat{\theta}_{\hat{M},T})}{\min \left\{ R(\theta, \hat{\theta}_{0,T}), R(\theta, \hat{\theta}_{1,T}) \right\}} \rightarrow 1, \quad \text{as } T \rightarrow \infty$$

Minimax-rate Adaptivity

$\hat{\theta}_{\hat{M},T}$ is **minimax-rate adaptive** if

$$\sup_T \left[\frac{\sup_{\theta \in \Theta} R(\theta, \hat{\theta}_{\hat{M},T})}{\inf_{\tilde{\theta}_T} \sup_{\theta \in \Theta} R(\theta, \tilde{\theta}_T)} \right] < \infty$$

The Strengths of AIC and BIC Cannot be Shared

Theorem

No model post-model selection estimator can be both pointwise-risk adaptive and minimax-rate adaptive.

AIC vs. BIC

- ▶ BIC is pointwise-risk adaptive but AIC is not. (This is effectively identical to consistency.)
- ▶ AIC is minimax-rate adaptive, but BIC is not.
- ▶ Further Reading: Yang (2005), Yang (2007)

Consistency and Efficiency in a Simple Example

Information Criteria

Consider criteria of the form $IC_m = 2\ell(\theta) - d_T \times \text{length}(\theta)$.

True DGP

$Y_1, \dots, Y_T \sim \text{iid } N(\mu, 1)$

Candidate Models

M_0 assumes $\mu = 0$, M_1 does not restrict μ . Only one parameter:

$$IC_0 = 2 \max_{\mu} \{\ell(\mu) : M_0\}$$

$$IC_1 = 2 \max_{\mu} \{\ell(\mu) : M_1\} - d_T$$

Log-Likelihood Function

Simple Algebra

$$\ell_T(\mu) = \text{Constant} - \frac{1}{2} \sum_{t=1}^T (Y_t - \mu)^2$$

Tedious Algebra

$$\sum_{t=1}^T (Y_t - \mu)^2 = T(\bar{Y} - \mu)^2 + T\hat{\sigma}^2$$

Combining These

$$\ell_T(\mu) = \text{Constant} - \frac{T}{2} (\bar{Y} - \mu)^2$$

The Selected Model \hat{M}

Information Criteria

M_0 sets $\mu = 0$ while M_1 uses the MLE \bar{Y} , so we have

$$IC_0 = 2 \max_{\mu} \{\ell(\mu) : M_0\} = 2 \times \text{Constant} - T\bar{Y}^2$$

$$IC_1 = 2 \max_{\mu} \{\ell(\mu) : M_1\} - d_T = 2 \times \text{Constant} - d_T$$

Difference of Criteria

$$IC_1 - IC_0 = T\bar{Y}^2 - d_T$$

Selected Model

$$\hat{M} = \begin{cases} M_1, & |\sqrt{T}\bar{Y}| \geq \sqrt{d_T} \\ M_0, & |\sqrt{T}\bar{Y}| < \sqrt{d_T} \end{cases}$$

Verifying Weak Consistency: $\mu \neq 0$

KL Divergence for M_0 and M_1

$$KL(g; M_0) = \mu^2/2, \quad KL(g; M_1) = 0$$

Condition on KL-Divergence

$$\liminf_{T \rightarrow \infty} \frac{1}{T} \sum_{t=1}^T \{KL(g; M_0) - KL(g; M_1)\} = \liminf_{n \rightarrow \infty} \frac{1}{T} \sum_{t=1}^T \left(\frac{\mu^2}{2} - 0 \right) > 0$$

Condition on Penalty

- ▶ Need $c_{T,k} = o_p(T)$, i.e. $c_{T,k}/T \xrightarrow{P} 0$.
- ▶ Both AIC and BIC satisfy this
- ▶ If $\mu \neq 0$, both AIC and BIC select M_1 wpa 1 as $T \rightarrow \infty$.

Verifying Consistency: $\mu = 0$

What's different?

- ▶ Both M_1 and M_0 are true and minimize KL divergence at zero.
- ▶ **Consistency** says choose most parsimonious true model: M_0

Verifying Conditions for Consistency

- ▶ $N(0, 1)$ model nested inside $N(\mu, 1)$ model
- ▶ Truth is $N(0, 1)$ so LR-stat is asymptotically $\chi^2(1) = O_p(1)$.
- ▶ For penalty term, need $\mathbb{P}(c_{T,k} - c_{T,0}) \rightarrow \infty$
- ▶ BIC satisfies this but AIC doesn't.

Finite-Sample Selection Probabilities: AIC

AIC Sets $d_T = 2$

$$\hat{M}_{AIC} = \begin{cases} M_1, & |\sqrt{T}\bar{Y}| \geq \sqrt{2} \\ M_0, & |\sqrt{T}\bar{Y}| < \sqrt{2} \end{cases}$$

$$\begin{aligned} P(\hat{M}_{AIC} = M_1) &= P(|\sqrt{T}\bar{Y}| \geq \sqrt{2}) \\ &= P(|\sqrt{T}\mu + Z| \geq \sqrt{2}) \\ &= P(\sqrt{T}\mu + Z \leq -\sqrt{2}) + [1 - P(\sqrt{T}\mu + Z \leq \sqrt{2})] \\ &= \Phi(-\sqrt{2} - \sqrt{T}\mu) + [1 - \Phi(\sqrt{2} - \sqrt{T}\mu)] \end{aligned}$$

where $Z \sim N(0, 1)$ since $\bar{Y} \sim N(\mu, 1/T)$ because $\text{Var}(Y_t) = 1$.

Finite-Sample Selection Probabilities: BIC

BIC sets $d_T = \log(T)$

$$\hat{M}_{BIC} = \begin{cases} M_1, & |\sqrt{T}\bar{Y}| \geq \sqrt{\log(T)} \\ M_0, & |\sqrt{T}\bar{Y}| < \sqrt{\log(T)} \end{cases}$$

Same steps as for the AIC except with $\sqrt{\log(T)}$ in the place of $\sqrt{2}$:

$$\begin{aligned} P(\hat{M}_{BIC} = M_1) &= P(|\sqrt{T}\bar{Y}| \geq \sqrt{\log(T)}) \\ &= \Phi(-\sqrt{\log(T)} - \sqrt{T}\mu) + [1 - \Phi(\sqrt{\log(T)} - \sqrt{T}\mu)] \end{aligned}$$

Interactive Demo: AIC vs BIC

https://fditraglia.shinyapps.io/CH_Figure_4_1/

Probability of Over-fitting

- ▶ If $\mu = 0$ both models are true but M_0 is more parsimonious.
- ▶ Probability of over-fitting (Z denotes standard normal):

$$\begin{aligned}P(\hat{M} = M_1) &= P(|\sqrt{T}\bar{Y}| \geq \sqrt{d_T}) = P(|Z| \geq \sqrt{d_T}) \\&= P(Z^2 \geq d_T) = P(\chi_1^2 \geq d_T)\end{aligned}$$

- ▶ AIC: $d_T = 2$ and $P(\chi_1^2 \geq 2) \approx 0.157$.
- ▶ BIC: $d_T = \log(T)$ and $P(\chi_1^2 \geq \log T) \rightarrow 0$ as $T \rightarrow \infty$.

AIC has $\approx 16\%$ prob. of over-fitting; BIC does not over-fit in the limit.

Risk of the Post-Selection Estimator

The Post-Selection Estimator

$$\hat{\mu} = \begin{cases} \bar{Y}, & |\sqrt{T}\bar{Y}| \geq \sqrt{d_T} \\ 0, & |\sqrt{T}\bar{Y}| < \sqrt{d_T} \end{cases}$$

Recall from above

Recall from above that $\sqrt{T}\bar{Y} = \sqrt{T}\mu + Z$ where $Z \sim N(0, 1)$

Risk Function

MSE risk times T to get risk relative to minimax rate: $1/T$.

$$R(\mu, \hat{\mu}) = T \cdot \mathbb{E} \left[(\hat{\mu} - \mu)^2 \right] = \mathbb{E} \left[\left(\sqrt{T}\hat{\mu} - \sqrt{T}\mu \right)^2 \right]$$

The Simplified MSE Risk Function

$$\begin{aligned}R(\mu, \hat{\mu}) &= 1 - [a\phi(a) - b\phi(b) + \Phi(b) - \Phi(a)] + T\mu^2 [\Phi(b) - \Phi(a)] \\&= 1 + [b\phi(b) - a\phi(a)] + (T\mu^2 - 1) [\Phi(b) - \Phi(a)]\end{aligned}$$

where

$$a = -\sqrt{d_T} - \sqrt{T}\mu$$

$$b = \sqrt{d_T} - \sqrt{T}\mu$$

https://fditraglia.shinyapps.io/CH_Figure_4_2/

Understanding the Risk Plot

AIC

- ▶ For any $\mu \neq 0$, risk $\rightarrow 1$ as $T \rightarrow \infty$, the risk of the MLE
- ▶ For $\mu = 0$, risk $\rightarrow 0$, risk of “zero” estimator
- ▶ Max risk is bounded

BIC

- ▶ For any $\mu \neq 0$, risk $\rightarrow 1$ as $T \rightarrow \infty$, the risk of the MLE
- ▶ For $\mu = 0$, risk $\rightarrow 0$, risk of “zero” estimator
- ▶ Max risk is unbounded

Lecture #5 – Andrews (1999) Moment Selection Criteria

Lighting Review of GMM

The J-test Statistic Under Correct Specification

The J-test Statistic Under Mis-specification

Andrews (1999; Econometrica)

Generalized Method of Moments (GMM) Estimation

Notation

Let v_t be a $(r \times 1)$ random vector, θ be a $(p \times 1)$ parameter vector, and f be a $(q \times 1)$ vector of real-valued functions.

Popn. Moment Conditions

$$\mathbb{E}[f(v_t, \theta_0)] = 0$$

Sample Moment Conditions

$$\bar{g}_T(\theta) = \frac{1}{T} \sum_{t=1}^T f(v_t, \theta)$$

GMM Estimator

$$\hat{\theta}_T = \arg \min_{\theta \in \Theta} \bar{g}_T(\theta)' \underset{(q \times q)}{W_T} \bar{g}_T(\theta), \quad W_T \rightarrow_p W \text{ (psd)}$$

Key Assumptions for GMM I

Stationarity

The sequence $\{v_t: -\infty < t < \infty\}$ is strictly stationary. This implies that *any* functions of v_t are constant over t .

Global Identification

$\mathbb{E}[f(v_t, \theta_0)] = 0$ but $\mathbb{E}[f(v_t, \tilde{\theta})] \neq 0$ for any $\tilde{\theta} \neq \theta_0$.

Regularity Conditions for Moment Functions

$f: \mathcal{V} \times \Theta \rightarrow \mathbb{R}^q$ satisfies:

- (i) f is v_t -almost surely continuous on Θ
- (ii) $E[f(v_t, \theta)] < \infty$ exists and is continuous on Θ

Key Assumptions for GMM I

Regularity Conditions for Derivative Matrix

- (i) $\nabla_{\theta'} f(v_t, \theta)$ exists and is v_t -almost continuous on Θ
- (ii) $E[\nabla_{\theta} f(v_t, \theta_0)] < \infty$ exists and is continuous in a neighborhood N_{ϵ} of θ_0
- (iii) $\sup_{\theta \in N_{\epsilon}} \left\| T^{-1} \sum_{t=1}^T \nabla_{\theta} f(v_t, \theta) - E[\nabla_{\theta} f(v_t, \theta)] \right\| \xrightarrow{P} 0$

Regularity Conditions for Variance of Moment Conditions

- (i) $E[f(v_t, \theta_0)f(v_t, \theta_0)']$ exists and is finite.
- (ii) $\lim_{T \rightarrow \infty} \text{Var} \left[\sqrt{T} \bar{g}_T(\theta_0) \right] = S$ exists and is a finite, positive definite matrix.

Main Results for GMM Estimation

Under the Assumptions Described Above

Consistency: $\hat{\theta}_T \xrightarrow{p} \theta_0$

Asymptotic Normality: $\sqrt{T}(\hat{\theta}_T - \theta_0) \xrightarrow{d} \mathcal{N}(0, MSM')$

$$M = (G_0'WG_0)^{-1}G_0'W$$

$$S = \lim_{T \rightarrow \infty} \text{Var} \left[\sqrt{T} \bar{g}_T(\theta_0) \right]$$

$$G_0 = E[\nabla_{\theta'} f(v_t, \theta_0)]$$

$$W = \text{plim}_{T \rightarrow \infty} W_T$$

The J-test Statistic

$$J_T = T \bar{g}_T(\hat{\theta}'_T) \hat{S}^{-1} \bar{g}_T(\hat{\theta}_T)$$

$$\hat{S} \rightarrow_p S = \lim_{T \rightarrow \infty} \text{Var} \left[\sqrt{T} \bar{g}_T(\theta_0) \right]$$

$$\bar{g}_T(\hat{\theta}_T) = \frac{1}{T} \sum_{t=1}^T f(v_t, \hat{\theta}_T)$$

$$\hat{\theta}_T = \text{GMM Estimator}$$

Case I: Correct Specification

Suppose that all of the preceding assumptions hold, in particular that the model is **correctly specified**:

$$\mathbb{E}[f(v_t, \theta_0)] = 0$$

Recall that under the standard assumptions, the GMM estimator is consistent **regardless of the choice of W_T** ...

Case I: Taylor Expansion under Correct Specification

$$W_T^{1/2} \sqrt{T} \bar{g}_T(\hat{\theta}_T) = [I_q - P(\theta_0)] W_T^{1/2} \sqrt{T} \bar{g}_T(\theta_0) + o_p(1)$$

$$P(\theta_0) = F(\theta_0) [F(\theta_0)' F(\theta_0)]^{-1} F(\theta_0)'$$

$$F(\theta_0) = W_T^{1/2} E[\nabla_{\theta} f(v_t, \theta_0)]$$

Over-identification

If $\dim(f) > \dim(\theta_0)$, $W_T^{1/2} \mathbb{E}[f(v_t, \theta_0)]$ is the linear combn. used in GMM estimation.

Identifying and Over-Identifying Restrictions

$P(\theta_0) \equiv$ **identifying restrictions**;

$I_q - P(\theta_0) \equiv$ **over-identifying restrictions**

J-test Statistic Under Correct Specification

$$W_T^{1/2} \sqrt{T} \bar{g}_T(\hat{\theta}_T) = [I_q - P(\theta_0)] W_T^{1/2} \sqrt{T} \bar{g}_T(\theta_0) + o_p(1)$$

- ▶ CLT for $\sqrt{T} \bar{g}_T(\theta_0)$
- ▶ $I_q - P(\theta_0)$ has rank $(q - p)$, since $P(\theta_0)$ has rank p .
- ▶ **Singular** normal distribution
- ▶ $W_T^{1/2} \sqrt{T} \bar{g}_T(\hat{\theta}_T) \xrightarrow{d} \mathcal{N}(0, N W_T^{1/2} S W_T^{1/2} N')$
- ▶ Substituting \hat{S}^{-1} , $J_T \xrightarrow{d} \chi_{q-p}^2$

Case II: Fixed Mis-specification

$$\mathbb{E}[f(v_t, \theta)] = \mu(\theta), \quad \|\mu(\theta)\| > 0, \quad \forall \theta \in \Theta$$

N.B.

This can *only* occur in the over-identified case, since we can always solve the population moment conditions in the just-identified case.

Notation

- ▶ $\theta^* \equiv$ solution to identifying restrictions ($\hat{\theta}_T \rightarrow_p \theta^*$)
- ▶ $\mu^* = \mu(\theta^*) = \text{plim}_{T \rightarrow \infty} \bar{g}_T(\hat{\theta}_T)$

Case II: Fixed Mis-specification

$$\frac{1}{T} J_T = \bar{g}_T(\hat{\theta}_T)' \hat{S}^{-1} \bar{g}_T(\hat{\theta}_T) = \mu_*' W \mu_* + o_p(1)$$

- ▶ W positive definite
- ▶ since $\mu(\theta) > 0$ for all $\theta \in \Theta$.
- ▶ Hence: $\mu_*' W \mu_* > 0$
- ▶ Fixed mis-specification $\Rightarrow J$ -test statistic *diverges at rate T* :

$$J_T = T \mu_*' W \mu_* + o_p(T)$$

Summary: Correct Specification vs. Fixed Mis-specification

Correct Specification: $J_T \Rightarrow \chi^2_{q-p} = O_p(1)$

Fixed Mis-specification: $J_T = O_p(T)$

Andrews (1999; Econometrica)

- ▶ Family of moment selection criteria (MSC) for GMM
- ▶ Aims to **consistently** choose any and all correct MCs and eliminate incorrect MCs
- ▶ AIC/BIC: add a **penalty** to maximized log-likelihood
- ▶ Andrews MSC: add a **bonus** term to the J-statistic
 - ▶ J-stat shows how well MCs “fit”
 - ▶ Compares $\hat{\theta}_T$ estimated using $P(\theta_0)$ to MCs from $I_q - P(\theta_0)$
 - ▶ J-stat tends to increase with degree of overidentification even if MCs are correct, since it converges to a χ^2_{q-p}

Andrews (1999) – Notation

$f_{max} \equiv (q \times 1)$ vector of all MCs under consideration

$c \equiv (q \times 1)$ selection vector: zeros and ones indicating which MCs are included

$\mathcal{C} \equiv$ set of all candidates c

$|c| \equiv \#$ of MCs in candidate c

Let $\hat{\theta}_T(c)$ be the efficient two-step GMM estimator based on the moment conditions $E[f(v_t, \theta, c)] = 0$ and define

$$V_{\theta}(c) = \left[G_0(c) S(c)^{-1} G_0(c) \right]^{-1}$$

$$G_0(c) = E[\nabla'_{\theta} f(v_t, \theta_0; c)]$$

$$S(c) = \lim_{T \rightarrow \infty} \text{Var} \left[\frac{1}{\sqrt{T}} \sum_{t=1}^T f(v_t, \theta_0; c) \right]$$

$$J_T(c) = T \bar{g}_T' \left(\hat{\theta}_T(c); c \right)' \hat{S}_T(c)^{-1} \bar{g}_T \left(\hat{\theta}_T(c); c \right)$$

Identification Condition

- ▶ Andrews wants maximal set of correct MCs
 - ▶ Consistent, minimum asymptotic variance
- ▶ But different θ values could solve $\mathbb{E}[f(v_t, \theta, c)]$ for different c !
- ▶ Which θ_0 are we actually trying to be consistent for?

More Notation

- ▶ $\mathcal{Z}^0 \equiv$ set of all c for which $\exists \theta$ with $\mathbb{E}[f(v_t, \theta, c)] = 0$
- ▶ $\mathcal{MZ}^0 \equiv$ subset of \mathcal{Z}^0 with **maximal** $|c|$.

Assumption

Andrews assumes that $\mathcal{MZ}^0 = \{c_0\}$, a singleton.

Family of Moment Selection Criteria

- ▶ Criteria of the form $MSC(c) = J_T(c) - B(T, |c|)$
- ▶ B is a **bonus term** that depends on sample size and # of MCs
- ▶ Choose $\hat{c}_T = \arg \min_{c \in \mathcal{C}} MSC(c)$
- ▶ Implementation Detail: Andrews suggests using a **centered** covariance matrix estimator:

$$\hat{S}(c) = \frac{1}{T} \sum_{t=1}^T \left[f(v_t, \hat{\theta}_T(c); c) - \bar{g}_T(\hat{\theta}_T(c); c) \right] \left[f(v_t, \hat{\theta}_T(c); c) - \bar{g}_T(\hat{\theta}_T(c); c) \right]'$$

based on the weighting matrix that *would be* efficient if the moment conditions were correctly specified. This remains consistent for $S(c)$ even under fixed mis-specification

Regularity Conditions for the J -test Statistic

- (i) If $\mathbb{E}[f(v_t, \theta; c)] = 0$ for a unique $\theta \in \Theta$, then $J_T(c) \xrightarrow{d} \chi^2_{|c|-p}$
- (ii) If $\mathbb{E}[f(v_t, \theta; c)] \neq 0$ for a *all* $\theta \in \Theta$ then $T^{-1}J_T(c) \xrightarrow{p} a(c)$, a finite, positive constant that may depend on c .

Regularity Conditions for Bonus Term

The bonus term can be written as $B(|c|, T) = \kappa_T h(|c|)$, where

- (i) $h(\cdot)$ is strictly increasing
- (ii) $\kappa_T \rightarrow \infty$ as $T \rightarrow \infty$ and $\kappa_T = o(T)$

Identification Conditions

- (i) $\mathcal{M}\mathcal{Z}^0 = \{c_0\}$
- (ii) $\mathbb{E}[f(v_t, \theta_0; c_0)] = 0$ and $E[f(v_t, \theta; c_0)] \neq 0$ for any $\theta \neq \theta_0$

Consistency of Moment Selection

Theorem

Under the preceding assumptions, $MSC(c)$ is a consistent moment selection criterion, i.e. $\hat{c}_T \xrightarrow{P} c_0$.

Some Examples

$$\text{GMM-BIC}(c) = J_T(c) - (|c| - p) \log(T)$$

$$\text{GMM-HQ}(c) = J_T(c) - 2.01 (|c| - p) \log(\log(T))$$

$$\text{GMM-AIC}(c) = J_T(c) - 2 (|c| - p) \log(T)$$

Corollary

GMM-AIC is *not* consistent, since $\kappa_T = 2$ does not diverge as $T \rightarrow \infty$.

GMM-BIC and GMM-HQ are consistent since $\lim_{T \rightarrow \infty} \log(T)/T = 0$ and $\lim_{T \rightarrow \infty} \log(\log(T))/T = 0$.