

Chapter 1

Maximum Likelihood Estimation

The word “estimation” is an euphemism of “educated guess”. What is a reasonable way to make a guess? The maximum likelihood principle is arguably one of the the most important ideas in the history of mathematical statistics. It is intuitive at the first glance, but it took centuries to formalize. Milestones were set by Sir Ronald Fisher (1890–1962) and his contemporaries. See the historical note by Stigler (2007).

In econometrics, the most familiar OLS estimator is indeed a special case of MLE when the error term is normally distribution. When the support of the dependence variable is not continuous, most estimators are constructed by assuming an underlying conditional density model and then estimated by MLE. LIML (limited information maximum likelihood) is 2SLS counterpart in the likelihood world.

1.1 Parametric Model

A parametric model is a complete specification of the distribution. Once the parameter is given, the distribution function is determined. Instead, a semiparametric model only gives a few features rather than a complete description of the distribution.

Example 1.1. Semiparametric model: If we know $Y \sim i.i.d. (\mu, \sigma^2)$, we can estimate μ, σ^2 by method of moments.

Parametric model: If we assume $Y \sim N(\mu, \sigma^2)$, the model has only two parameters μ and σ^2 .

Definition 1.1. Parametric model. The distribution of the data $\mathbf{Y} = (Y_1, \dots, Y_n)$ is known up to a finite dimensional parameter.

Let Θ be the parameter space a researcher specifies.

Definition 1.2. A model is **correctly specified**, if the true DGP is $f(\mathbf{Y} | \theta_0)$ for some $\theta_0 \in \Theta$. Otherwise, the model is **misspecified**.

1.2 Likelihood

In this chapter we will mostly talk about unconditional models. The results can be carried over to conditional models. To keep the setting simple, let $\mathbf{Y} = (Y_1, \dots, Y_n)$ be i.i.d. The **likelihood** of the sample under a hypothesized value of $\theta \in \Theta$ is is

$$L(\theta; \mathbf{Y}) = f(\mathbf{Y}; \theta) = \prod_{i=1}^n f(Y_i; \theta_0).$$

Remark 1.1. In the above expression $L(\theta; \mathbf{Y})$ and $f(\mathbf{Y}; \theta)$ are equal, but they have different interpretations. The former emphasizes that when the data \mathbf{Y} is provided L is a function of the parameter θ . The latter is the opposite: given the parameter θ , the joint density is a function of the realized value \mathbf{Y} .

Remark 1.2. In this note I adhere to the frequentist view where θ is an unknown constant, and I thus prefer using “;” to separate the two arguments. However, some textbooks (including Hansen’s) write $L(\theta|\mathbf{Y})$ and $f(\mathbf{Y}|\theta)$, which I don’t like because “|” delivers a confusing connotation of conditional density (for example $f(Y|X)$) as if θ is treated as a random variable. I would adopt “|” only in the Bayesian world, where the posterior density of the random θ is updated via the Bayes Theorem

$$L(\theta|\mathbf{Y}) = \frac{f(\mathbf{Y}|\theta) \pi(\theta)}{f(\mathbf{Y})},$$

give the prior distribution $\pi(\theta)$.

Example 1.2. Conditional model: the conditioning variable can be viewed as if it is fixed and the randomness comes from the error term only.

$$Y_i = X_i' \beta + \varepsilon_i$$

X_i is the conditional variable. The condition $E(\varepsilon_i | X_i) = 0$ together with a full rank $E[X_i X_i']$ can help to identify β . This is semiparametric model. However, if we assume $f(\varepsilon_i | X_i) \sim N(0, \sigma^2)$, then conditional parametric model as it completely describes $f(Y_i | X_i)$ and it becomes a conditional parametric model with likelihood $L((\beta, \sigma^2); \mathbf{Y}|\mathbf{X})$.

The **log-likelihood** is

$$\ell_n(\theta) = \frac{1}{n} \log L(\theta; \mathbf{Y}) = \frac{1}{n} \sum_{i=1}^n \log f(Y_i; \theta).$$

Here, we put $1/n$ to average the log-likelihood. This scaling factor does not change the estimation at all.

In practice, we work with the log-likelihood, which is more convenient. the MLE estimator is defined as

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

To justify the likelihood principle, consider the population version of the

$$\ell(\theta) = E[\log f(Y; \theta)]$$

Theorem 1.1. When model is correctly specified, θ_0 is the maximizer.

Proof. The Kullback-Leibler distance

$$\begin{aligned} E[\log p(\theta_0)] - E[\log p(\theta)] &= E[\log(p(\theta_0)/p(\theta))] \\ &= -E[\log(p(\theta)/p(\theta_0))] \\ &\geq -\log E[p(\theta)/p(\theta_0)] = 0 \end{aligned}$$

where the inequality holds by Jensen’s inequality for the convex function $-\log(\cdot)$. □

1.3 Score, Hessian, and Information

Define the **score** as

$$\psi_n(\theta) = \sum_{i=1}^n \frac{\partial}{\partial \theta} \log f(Y_i; \theta)$$

and the **Hessian** as

$$\mathcal{H}_n(\theta) = - \sum_{i=1}^n \frac{\partial^2}{\partial \theta \partial \theta'} \log f(Y_i; \theta).$$

Moreover, define the **efficient score**

$$\psi_0 = \frac{\partial}{\partial \theta} \log f(Y_i; \theta_0)$$

as the first derivative of the log-likelihood of a representative observation Y_i evaluated at θ_0 .

Theorem 1.2. *If the model is correctly specified, the support of Y does not depend on θ , and θ_0 is in the interior of Θ , then $E(\psi_0) = 0$.*

Proof. By the Leibniz integral rule,¹

$$\begin{aligned} E(\psi(\theta)) &= E \left[\frac{\partial}{\partial \theta} \log f(Y_i; \theta) \right] = \int \frac{\partial}{\partial \theta} \log f(Y_i; \theta) dF_0(x) \\ &= \frac{\partial}{\partial \theta} \int \log f(Y_i; \theta) dF_0(x) = \frac{\partial}{\partial \theta} E[\log f(Y_i; \theta)]. \end{aligned}$$

Evaluate $\frac{\partial}{\partial \theta} E[\log f(Y_i; \theta)]$ as θ_0 , we have $\frac{\partial}{\partial \theta} E[\log f(Y_i; \theta_0)] = 0$ as $E[\log f(Y_i; \theta_0)]$ is maximized at θ_0 , which is an interior. \square

The **Fisher information matrix**

$$\mathcal{I}_0 = E[\psi_0 \psi_0']$$

is the variance of the efficient score. The **expected Hessian** is

$$\mathcal{H}_0 = -E \left[\frac{\partial^2}{\partial \theta \partial \theta'} \log f(Y; \theta_0) \right].$$

Since $E \left[\frac{\partial^2}{\partial \theta \partial \theta'} \log f(Y; \theta_0) \right]$ is negative definite, in this definition the expected Hessian is positive-definite.

Theorem 1.3. *If the model is correctly specified, we have the **information matrix equality**: $\mathcal{I}_0 = \mathcal{H}_0$.*²

¹Leibniz integral rule (from undergraduate calculus): For an integral of the form $\int_{a(t)}^{b(t)} g(t, z) dz$, the derivative

$$\frac{d}{dt} \int_{a(t)}^{b(t)} g(t, z) dz = g(t, b(t)) \cdot \frac{d}{dt} b(t) - g(t, a(t)) \cdot \frac{d}{dt} a(t) + \int_{a(t)}^{b(t)} \frac{\partial}{\partial t} g(t, z) dz.$$

²Some textbooks may define the expected Hessian as $\mathcal{H}_0 = E \left[\frac{\partial^2}{\partial \theta \partial \theta'} \log f(Y; \theta_0) \right]$. This is also valid if we rewrite the information matrix equality as $\mathcal{I}_0 + \mathcal{H}_0 = 0$.

Proof. Start with Hessian,

$$\begin{aligned} E \left[\frac{\partial^2}{\partial \theta \partial \theta'} \log f(\theta_0) \right] &= E \left[\frac{\partial}{\partial \theta} \frac{\partial}{\partial \theta'} \log f(\theta_0) \right] = E \left[\frac{\partial}{\partial \theta} \frac{\frac{\partial}{\partial \theta'} f(\theta)}{f(\theta)} \Big|_{\theta=\theta_0} \right] \\ &= -E \left[\frac{\frac{\partial}{\partial \theta} f(\theta) \frac{\partial}{\partial \theta'} f(\theta)}{f^2(\theta)} \Big|_{\theta=\theta_0} \right] + E \left[\frac{\frac{\partial^2}{\partial \theta \partial \theta'} f(\theta)}{f(\theta)} \Big|_{\theta=\theta_0} \right]. \end{aligned}$$

The first term:

$$E \left[\frac{\frac{\partial}{\partial \theta} f(\theta) \frac{\partial}{\partial \theta'} f(\theta)}{f^2(\theta)} \right] = E \left[\frac{\partial}{\partial \theta} \log f(\theta) \frac{\partial}{\partial \theta'} \log f(\theta) \right] = E [\psi_0 \psi_0'] = \mathcal{J}_0.$$

The second term: $E \left[\frac{\frac{\partial^2}{\partial \theta \partial \theta'} f(\theta)}{f(\theta)} \right] = \int \frac{\frac{\partial^2}{\partial \theta \partial \theta'} f(\theta)}{f(\theta)} f(\theta) dx$. If the model is correctly specified, evaluated at $\theta = \theta_0$ the $f(\theta_0)$ cancels out, leading to

$$\int \frac{\frac{\partial^2}{\partial \theta \partial \theta'} f(\theta_0)}{f(\theta_0)} f(\theta_0) dx = \int \frac{\partial^2}{\partial \theta \partial \theta'} f(\theta_0) dx = \frac{\partial^2}{\partial \theta \partial \theta'} \int f(\theta_0) dx = \frac{\partial^2}{\partial \theta \partial \theta'} 1 = 0.$$

We complete the proof. \square

Notice that the information matrix equality holds only when the model is correctly specified. It fails when the model is misspecified.

1.4 Cramér-Rao Lower Bound

Theorem 1.4. Suppose the model is correctly specified, the support of Y does not depend on θ , and θ_0 is in the interior of Θ . If $\tilde{\theta}$ is unbiased estimator, then $\text{var}(\tilde{\theta}) \geq (n\mathcal{J}_0)^{-1}$.

Proof. Because of unbiasedness,

$$\theta = E_{\theta} [\tilde{\theta}] = \int \tilde{\theta} f(\mathbf{Y}; \theta) d\mathbf{y}$$

for any $\theta \in \Theta$. \mathbf{Y} here is for the entire sample, $f(\mathbf{Y}; \theta) = f(Y_1, \dots, Y_n; \theta) = \prod_{i=1}^n f(Y_i; \theta)$. Take derivative at the two sides. The LHS is

$$\frac{\partial \theta}{\partial \theta'} = \mathbf{I}_p$$

. The RHS:

$$\begin{aligned} \frac{\partial}{\partial \theta'} \int \tilde{\theta} f(\mathbf{Y}; \theta) d\mathbf{y} &= \int \tilde{\theta} \frac{\partial}{\partial \theta'} f(\mathbf{Y}; \theta) d\mathbf{y} \\ &= \int \tilde{\theta} \frac{\frac{\partial}{\partial \theta'} f(\mathbf{Y}; \theta)}{f(\mathbf{Y}; \theta)} f(\mathbf{Y}; \theta) d\mathbf{y} \\ &= \int \tilde{\theta} \frac{\partial}{\partial \theta'} \log f(\mathbf{Y}; \theta) f(\mathbf{Y}; \theta) d\mathbf{y} \\ &= \int \tilde{\theta} \psi_n(\theta) f(\mathbf{Y}; \theta) d\mathbf{y} \end{aligned}$$

Evaluate at the true θ_0 , and due to i.i.d. data

$$\mathbf{I}_p = \int \tilde{\theta} \psi_n(\theta_0) f(\mathbf{Y}; \theta_0) d\mathbf{Y} = E[\tilde{\theta} \psi_n(\theta_0)] = E[(\tilde{\theta} - \theta_0) \psi_n(\theta_0)]$$

where the last equality holds by $E[\theta_0 \psi_n(\theta_0)] = \theta_0 E[\psi_n(\theta_0)] = \theta_0 E[n\psi_0] = 0$. We thus have

$$\text{var} \begin{pmatrix} \tilde{\theta} - \theta_0 \\ \psi_n(\theta_0) \end{pmatrix} = \begin{bmatrix} \mathbf{V} & \mathbf{I}_p \\ \mathbf{I}_p & n\mathcal{J}_0 \end{bmatrix}.$$

Pre- and post-multiply $\begin{bmatrix} \mathbf{I}_p & -(n\mathcal{J}_0)^{-1} \end{bmatrix}$, we have

$$\begin{bmatrix} \mathbf{I}_p & -(n\mathcal{J}_0)^{-1} \end{bmatrix} \begin{bmatrix} \mathbf{V} & \mathbf{I}_p \\ \mathbf{I}_p & n\mathcal{J}_0 \end{bmatrix} \begin{bmatrix} -\mathbf{I}_p \\ -(n\mathcal{J}_0)^{-1} \end{bmatrix} = \mathbf{V} - (n\mathcal{J}_0)^{-1} \geq 0.$$

□

The Cramér-Rao Lower Bound is a lower bound. It may not be reachable. When it is reached, an estimator is **Cramér-Rao efficient** if it is unbiased and the variance is $(n\mathcal{J}_0)^{-1}$.

Example 1.3. Normal distribution: Let $\gamma = \sigma^2$

$$\log \ell_n(Y; \mu, \sigma^2) = -\frac{n}{2} \log \gamma - \frac{n}{2} \log \pi - \frac{1}{2\gamma} \sum_{i=1}^n (Y_i - \mu)^2$$

$$\psi_n(\mu, \sigma^2) = \begin{cases} \frac{1}{\gamma} \sum_{i=1}^n (Y_i - \mu) \\ -\frac{n}{2\gamma} + \frac{1}{2\gamma^2} \sum_{i=1}^n (Y_i - \mu)^2 \end{cases}$$

$$\mathcal{H}_n(\mu, \sigma^2) = \begin{bmatrix} \frac{n}{\gamma} & \frac{1}{2\gamma^2} \sum_{i=1}^n (Y_i - \mu) \\ \frac{1}{2\gamma^2} \sum_{i=1}^n (Y_i - \mu) & -\frac{n}{2\gamma^2} + \frac{1}{\gamma^3} \sum_{i=1}^n (Y_i - \mu)^2 \end{bmatrix}$$

Expected Hessian:

$$E[\mathcal{H}_n(\mu, \sigma^2)] = \begin{bmatrix} \frac{n}{\gamma} & 0 \\ 0 & \frac{n}{2\gamma^2} \end{bmatrix}$$

Take inverse:

$$\begin{bmatrix} \frac{\gamma}{n} & 0 \\ 0 & 2\frac{\gamma^2}{n} \end{bmatrix}$$

This is the lower bound.

Check:

the sample mean:

$$\text{var} \left(\frac{1}{n} \sum_{i=1}^n Y_i \right) = \frac{\sigma^2}{n}$$

The sample mean is Cramér-Rao efficient.

$$s_n^2 = \frac{1}{n-1} \sum_{i=1}^n (Y_i - \bar{Y})^2 = \frac{1}{n-1} Y' \left(I - \frac{1}{n} \mathbf{1}_n \mathbf{1}_n' \right) Y$$

$E(S_n^2) = \sigma^2$ is unbiased

$$(n-1) \frac{s_n^2}{\sigma^2} = \left(\frac{Y}{\sigma} \right)' \left(I - \frac{1}{n} \mathbf{1}_n \mathbf{1}_n' \right) \left(\frac{Y}{\sigma} \right) \sim \chi^2(n-1)$$

So,

$$s_n^2 = \frac{\chi^2(n-1)}{n-1} \sigma^2$$

$$\text{var}(s_n^2) = \frac{\sigma^4}{(n-1)^2} 2(n-1) = \frac{2\sigma^4}{n-1} > \frac{2\sigma^4}{n}$$

Does not satisfy Cramér-Rao efficient.

1.5 Asymptotic Normality

MLE is a special case of m-estimator. Under regularity conditions, $\hat{\theta} \xrightarrow{p} \theta_0$, and asymptotically normal:

$$\sqrt{n}(\hat{\theta} - \theta_0) \xrightarrow{d} N(0, \mathcal{H}_0^{-1} \mathcal{J}_0 \mathcal{H}_0^{-1})$$

When the information equality holds, the asymptotic variance is simplified as $\mathcal{J}_0^{-1} \mathcal{J}_0 \mathcal{J}_0^{-1} = \mathcal{J}_0^{-1}$, and thus

$$\sqrt{n}(\hat{\theta} - \theta_0) \xrightarrow{d} N(0, \mathcal{J}_0^{-1}).$$

In other words, it achieves asymptotic efficiency.

Caveat:

1. need correct specification
2. the comparison is restricted to asymptotically unbiased estimator. There are biased estimators with better overall performance.

1.6 Kullback-Leibler Divergence

$$KLIC(f, g) = \int f(x) \log \frac{f(x)}{g(x)} dx$$

Properties:

1. $KLIC(f, f) = 0$
2. $KLIC(f, g) \geq 0$
3. $f = \arg \min_g KLIC(f, g)$

If $f(x) = f(x; \theta)$ is a parametric family

$$\theta_0 = \arg \min_{\theta \in \Theta} KLIC(f, f_\theta)$$

which is correctly specified model.

Pseudo-true parameter:

$$\theta_0 = \arg \min_{\theta \in \Theta} KLIC(f, f_\theta)$$

which is misspecified model.

KLIC is the distance measure of any two distributions.

$$\begin{aligned} KLIC(f, f_\theta) &= \int f(x) \log f(x) dx - \int f(x) \log f(x; \theta) dx \\ &= \int f(x) \log f(x) dx - E[\log f(x; \theta)] \\ &= \int f(x) \log f(x) dx - \ell(\theta) \end{aligned}$$

the pseudo-true value

$$\theta^* = \arg \max_{\theta \in \Theta} \ell(\theta)$$

The information equality was proved under correct specification. When the model is misspecified,

$$E[S(\theta^*) S(\theta^*)'] \neq E\left[\frac{\partial^2}{\partial \theta \partial \theta'} \log f(\theta^*)\right].$$

As a result, we will have a sandwich-form asymptotic variance in

$$\sqrt{n}(\hat{\theta} - \theta^*) \xrightarrow{d} N(0, \mathcal{H}_*^{-1} \mathcal{I}_* \mathcal{H}_*^{-1})$$

understand that \mathcal{I}_* and \mathcal{H}_* are evaluated at the pseudo-true value.

Zhentao Shi. January 24, 2024

Bibliography

Stigler, S. M. (2007). The epic story of maximum likelihood. *Statistical Science*, 598–620. 1