

Least Squares

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Notation: y_i is a scalar, and x_i is a $K \times 1$ vector. Y is an $n \times 1$ vector, and X is an $n \times K$ matrix.

1 Algebra of Least Squares

1.1 OLS estimator

As we have learned from the linear project model, the projection coefficient β in the regression

$$y_i = x_i' \beta + e_i$$

can be written as $\beta = (E[x_i x_i'])^{-1} E[x_i y_i]$. While population is something imaginary, in reality we possess a sample of n observations $(y_i, x_i)_{i=1}^n$. We thus replace the population mean $E[\cdot]$ by the sample mean, and the result-

ing estimator is

$$\hat{\beta} = \left(\frac{1}{n} \sum_{i=1}^n x_i x_i' \right)^{-1} \frac{1}{n} \sum_{i=1}^n x_i y_i = (X'X)^{-1} X'y.$$

This is one way to motivate the OLS estimator.

Alternatively, we can derive the OLS estimator from minimizing the sum of squared residuals

$$Q(\beta) = \sum_{i=1}^n (y_i - x_i' \beta)^2 = (Y - X\beta)' (Y - X\beta).$$

Solve the first-order condition

$$\frac{\partial}{\partial \beta} Q(\beta) = -2X'(Y - X\beta) = 0.$$

This necessary condition for optimality gives exactly the same $\hat{\beta}$. Moreover, the second-order condition

$$\frac{\partial^2}{\partial \beta \partial \beta'} Q(\beta) = 2X'X$$

shows that $Q(\beta)$ is convex in β due to the positive semi-definite matrix $X'X$. ($Q(\beta)$ is strictly convex in β if $X'X$ is positive definite.)

Here are some definitions and properties of the OLS estimator.

- Fitted value: $\hat{Y} = X\hat{\beta}$.
- Projector: $P_X = X(X'X)^{-1}X'$; Annihilator: $M_X = I_n - P_X$.

- $P_X M_X = M_X P_X = 0$.
- If $AA = A$, we call it an idempotent matrix. Both P_X and M_X are idempotent.
- Residual: $\hat{e} = Y - \hat{Y} = Y - X\hat{\beta} = Y - X(X'X)^{-1}X'Y = (I - P_X)Y = M_X Y = M_X(X\beta + e) = M_X e$. (Note: $M_X X = (I - P_X)X = X - X = 0 \implies M_X X\beta = 0$)
- $X'\hat{e} = X'M_X e = 0$. (Note again $X'M_X = 0$)
- $\frac{1}{n} \sum_{i=1}^n \hat{e}_i = 0$ if x_i contains a constant.

(Justification: $X'\hat{e} = \begin{bmatrix} 1 & 1 & \cdots & 1 \\ * & * & \cdots & * \\ \cdots & \cdots & \ddots & \vdots \\ * & * & \cdots & * \end{bmatrix} \begin{bmatrix} \hat{e}_1 \\ \hat{e}_2 \\ \vdots \\ \hat{e}_n \end{bmatrix} = \begin{bmatrix} 0 \\ 0 \\ \vdots \\ 0 \end{bmatrix}$ and the first row implies $\sum_{i=1}^n \hat{e}_i = 0$.)

1.2 Goodness of Fit

The so-called *R-squared* is a popular measure of goodness-of-fit in the linear regression. R-squared is well defined only when a constant is included in the regressors. Let $M_\iota = I_n - \frac{1}{n}\iota\iota'$, where ι is an $n \times 1$ vector of 1's. M_ι is the *demeaner*, in the sense that $M_\iota(z_1, \dots, z_n)' = (z_1 - \bar{z}, \dots, z_n - \bar{z})'$, where $\bar{z} = \frac{1}{n} \sum_{i=1}^n z_i$. For any X , we can decompose $Y = P_X Y + M_X Y =$

$\hat{Y} + \hat{e}$. The total variation is

$$Y' M_l Y = (\hat{Y} + \hat{e})' M_l (\hat{Y} + \hat{e}) = \hat{Y}' M_l \hat{Y} + 2\hat{Y}' M_l \hat{e} + \hat{e}' M_l \hat{e} = \hat{Y}' M_l \hat{Y} + \hat{e}' \hat{e}$$

where the last equality follows by $M_l \hat{e} = \hat{e}$ as $\frac{1}{n} \sum_{i=1}^n \hat{e}_i = 0$, and $\hat{Y}' \hat{e} = Y' P_X M_X e = 0$. R-squared is defined as $\hat{Y}' M_l \hat{Y} / Y' M_l Y$.

1.3 Frish-Waugh-Lovell Theorem

The Frish-Waugh-Lovell (FWL) theorem is an algebraic fact about the formula of a subvector of the OLS estimator. To derive the FWL theorem We need to use the inverse of partitioned matrix. For a positive definite

symmetric matrix $A = \begin{pmatrix} A_{11} & A_{12} \\ A_{12}' & A_{22} \end{pmatrix}$, the inverse can be written as

$$A^{-1} = \begin{pmatrix} (A_{11} - A_{12} A_{22}^{-1} A_{12}')^{-1} & - (A_{11} - A_{12} A_{22}^{-1} A_{12}')^{-1} A_{12} A_{22}^{-1} \\ - A_{22}^{-1} A_{12}' (A_{11} - A_{12} A_{22}^{-1} A_{12}')^{-1} & (A_{22} - A_{12}' A_{11}^{-1} A_{12})^{-1} \end{pmatrix}.$$

In our context of OLS estimator, let $X = \begin{pmatrix} X_1 & X_2 \end{pmatrix}$

$$\begin{aligned}
\hat{\beta} &= \begin{pmatrix} \hat{\beta}_1 \\ \hat{\beta}_2 \end{pmatrix} = (X'X)^{-1}X'Y \\
&= \left(\begin{pmatrix} X_1' \\ X_2' \end{pmatrix} \begin{pmatrix} X_1 & X_2 \end{pmatrix} \right)^{-1} \begin{pmatrix} X_1'Y \\ X_2'Y \end{pmatrix} \\
&= \begin{pmatrix} X_1'X_1 & X_1'X_2 \\ X_2'X_1 & X_2'X_2 \end{pmatrix}^{-1} \begin{pmatrix} X_1'Y \\ X_2'Y \end{pmatrix} \\
&= \begin{pmatrix} (X_1'M'_{X_2}X_1)^{-1} & - (X_1'M'_{X_2}X_1)^{-1} X_1'X_2 (X_2'X_2)^{-1} \\ * & * \end{pmatrix} \begin{pmatrix} X_1'Y \\ X_2'Y \end{pmatrix}.
\end{aligned}$$

The subvector

$$\begin{aligned}
\hat{\beta}_1 &= (X_1'M'_{X_2}X_1)^{-1} X_1'Y - (X_1'M'_{X_2}X_1)^{-1} X_1'X_2 (X_2'X_2)^{-1} X_2'Y \\
&= (X_1'M'_{X_2}X_1)^{-1} (X_1'Y - X_1'P_{X_2}Y) \\
&= (X_1'M'_{X_2}X_1)^{-1} X_1'M_{X_2}Y.
\end{aligned}$$

Notice that $\hat{\beta}_1$ can be obtained by the following:

1. Regress y on X_2 , obtain residuals \tilde{e}_2 ;
2. Regress X_1 on X_2 , obtain residuals \tilde{X}_2 ;
3. Regress \tilde{e}_2 on \tilde{X}_2 , obtain OLS estimates $\hat{\beta}_1$.

Similar derivation can also be carried out in the population linear projection. See Hansen (2019) Chapter 2.22-23.

2 Statistical Properties of Least Squares

In this section we return to the classical statistical framework under restrictive distributional assumption $y_i|x_i \sim N(x_i'\beta, \gamma)$, where $\gamma = \sigma^2$.

2.1 Maximum Likelihood Estimation

The *conditional* likelihood of observing a *random sample* $(y_i, x_i)_{i=1}^n$ is

$$\prod_{i=1}^n \frac{1}{\sqrt{2\pi\gamma}} \exp\left(-\frac{1}{2\gamma} (y_i - x_i'\beta)^2\right),$$

and the (conditional) log-likelihood function is

$$L(\beta, \gamma) = -\frac{n}{2} \log 2\pi - \frac{n}{2} \log \gamma - \frac{1}{2\gamma} \sum_{i=1}^n (y_i - x_i'\beta)^2.$$

The maximum likelihood estimator $\hat{\beta}_{MLE}$ can be found using the FOC:

$$\frac{\partial}{\partial \beta} L(\beta, \gamma) = \frac{1}{2\gamma} \sum_{i=1}^n 2x_i (y_i - x_i'\beta) = 0.$$

Rearranging the above equation in matrix form $X'Y = X'X\hat{\beta}_{MLE}$, we explicitly solve

$$\hat{\beta}_{MLE} = (X'X)^{-1}X'Y.$$

The maximum likelihood estimator (MLE) coincides with the OLS estimator. Similarly, another FOC gives $\hat{\gamma}_{\text{MLE}} = \hat{e}'\hat{e}/n$.

2.2 Classical Finite Sample Distribution

We can show the finite-sample exact distribution of $\hat{\beta}$ assuming the error term follows a Gaussian distribution. *Finite sample distribution* means that the distribution holds for any n ; it is in contrast to *asymptotic distribution*, which is a large sample approximation to the finite sample distribution. Let the “error term” $e_i = y_i - x_i'\beta$, we have $e_i|x_i = y_i|x_i - x_i'\beta \sim N(0, \gamma)$. Since the conditional distribution of e_i on x_i is invariant with x_i , the discrepancy e_i is statistical independent of x_i . Assume The estimator

$$\hat{\beta} = (X'X)^{-1} X'Y = (X'X)^{-1} X' (X'\beta + e) = \beta + (X'X)^{-1} X'e,$$

and its conditional distribution can be written as

$$\begin{aligned} \hat{\beta}|X &= \beta + (X'X)^{-1} X'e|X \\ &\sim \beta + (X'X)^{-1} X' \cdot N(0_n, \sigma^2 \cdot I_n) \\ &\sim N(\beta, \sigma^2 (X'X)^{-1} X'X (X'X)^{-1}) \sim N(\beta, \sigma^2 (X'X)^{-1}). \end{aligned}$$

The k -th element of the vector coefficient

$$\hat{\beta}_k|X = \eta_k' \hat{\beta}|X \sim N(\beta_k, \sigma^2 \eta_k' (X'X)^{-1} \eta_k) \sim N(\beta_k, \sigma^2 (X'X)^{-1}_{kk}),$$

where $\eta_k = (1 \{l = k\})_{l=1,\dots,K}$ is the selector of the k -th element.

In reality, σ^2 is an unknown parameter, and

$$s^2 = \hat{e}'\hat{e} / (n - K) = e' M_X e / (n - K)$$

is an unbiased estimator of σ^2 . Consider the t -statistic

$$\begin{aligned} T_k &= \frac{\hat{\beta}_k - \beta_k}{\sqrt{s^2 [(X'X)^{-1}]_{kk}}} \\ &= \frac{\hat{\beta}_k - \beta_k}{\sqrt{\sigma^2 [(X'X)^{-1}]_{kk}}} \cdot \frac{\sqrt{\sigma^2}}{\sqrt{s^2}} \\ &= \frac{(\hat{\beta}_k - \beta_k) / \sqrt{\sigma^2 [(X'X)^{-1}]_{kk}}}{\sqrt{\frac{e'}{\sigma} M_X \frac{e}{\sigma} / (n - K)}}. \end{aligned}$$

The numerator follows a standard normal, and the denominator follows $\frac{1}{n-K} \chi^2 (n - K)$. Moreover, the numerator and the denominator are statistically independent (Basu's theorem). As a result, we conclude $T_k \sim t (n - K)$. This finite sample distribution is crucial when conducting statistical inference.

2.3 Mean and Variance

Now we relax the normality assumption and statistical independence. Instead, we represent the regression model as $y_i = x_i' \beta + e_i$ and

$$\begin{aligned} E[e|X] &= 0 \\ \text{var}[e|X] &= \sigma^2 I_n. \end{aligned}$$

where the first condition is the *mean independence* assumption, and the second condition is the *homoskedasticity* assumption. These assumptions are about the first and second *moments* of e_i conditional on x_i . Unlike the normality assumption, they do not restrict the *distribution* of e_i .

- Unbiasedness:

$$E[\hat{\beta}|X] = E[(X'X)^{-1} X'Y|X] = E[(X'X)^{-1} X'(X'\beta + e)|X] = \beta.$$

Unbiasedness does not rely on homoskedasticity.

- Variance:

$$\begin{aligned}
\text{var}(\hat{\beta}|X) &= E \left[(\hat{\beta} - E\hat{\beta}) (\hat{\beta} - E\hat{\beta})' | X \right] \\
&= E \left[(\hat{\beta} - \beta) (\hat{\beta} - \beta)' | X \right] \\
&= E \left[(X'X)^{-1} X' e e' X (X'X)^{-1} | X \right] \\
&= (X'X)^{-1} X' E [e e' | X] X (X'X)^{-1} \\
&= (X'X)^{-1} X' (\sigma^2 I_n) X (X'X)^{-1} \\
&= \sigma^2 (X'X)^{-1}.
\end{aligned}$$

Homoskedasticity is essential in this derivation.

Example (Heteroskedasticity) If $e_i = x_i u_i$, where x_i is a scalar random variable, u_i is independent of x_i , $E[u_i] = 0$ and $E[u_i^2] = \sigma^2$. Then $E[e_i|x_i] = 0$ but $E[e_i^2|x_i] = \sigma^2 x_i^2$ is a function of x_i . We say e_i^2 is a heteroskedastic error.

2.4 Gauss-Markov Theorem

Gauss-Markov theorem justifies the OLS estimator as the efficient estimator among all linear unbiased ones. *Efficient* here means that it enjoys the smallest variance in a family of estimators.

There are numerous linearly unbiased estimators. For example, $(Z'X)^{-1} Z'y$ for $z_i = x_i^2$ is unbiased because $E \left[(Z'X)^{-1} Z'y \right] = E \left[(Z'X)^{-1} Z' (X\beta + e) \right] = \beta$.

Let $\tilde{\beta} = A'y$ be a generic linear estimator, where A is any $n \times K$ func-

tions of X . As

$$E [A'y|X] = E [A' (X\beta + e) |X] = A'X\beta.$$

So the linearity and unbiasedness of $\tilde{\beta}$ implies $A'X = I_n$. Moreover, the variance

$$\text{var} (A'y|X) = E [(A'y - \beta) (A'y - \beta)' |X] = E [A'ee'A|X] = \sigma^2 A'A.$$

Let $C = A - X (X'X)^{-1}$.

$$\begin{aligned} A'A - (X'X)^{-1} &= (C + X (X'X)^{-1})' (C + X (X'X)^{-1}) - (X'X)^{-1} \\ &= C'C + (X'X)^{-1} X'C + C'X (X'X)^{-1} \\ &= C'C, \end{aligned}$$

where the last equality follows as

$$(X'X)^{-1} X'C = (X'X)^{-1} X' (A - X (X'X)^{-1}) = (X'X)^{-1} - (X'X)^{-1} = 0.$$

Therefore $A'A - (X'X)^{-1}$ is a positive semi-definite matrix. The variance of any $\tilde{\beta}$ is no smaller than the OLS estimator $\hat{\beta}$.

Homoskedasticity is a restrictive assumption. Under homoskedasticity, $\text{var} (\hat{\beta}) = \sigma^2 (X'X)^{-1}$. Popular estimator of σ^2 is the sample mean of the residuals $\hat{\sigma}^2 = \frac{1}{n} \hat{e}'\hat{e}$ or the unbiased one $s^2 = \frac{1}{n-K} \hat{e}'\hat{e}$. Under het-

eroskedasticity, Gauss-Markov theorem does not apply.