Classical linear model [*]

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1.1. Notations

We wish to explain or predict a variable y through k other $x_1, x_2, \ldots, x_k$. We T observations on each variable:

$$
\begin{aligned}
y &= \begin{pmatrix} y_1 \\ y_2 \\ \vdots \\ y_T \end{pmatrix} \text{: dependent variable (to explain)} \\
x_i &= \begin{pmatrix} x_{1i} \\ x_{2i} \\ \vdots \\ x_{Ti} \end{pmatrix}, \quad i = 1, \ldots, k \text{: explanatory variables.}
\end{aligned}
$$

Usually, the explanatory variables are represented by the $T \times k$ matrix

$$
X = [x_1, x_2, \ldots, x_k] = \begin{bmatrix} x_{11} & x_{12} & \cdots & x_{1k} \\ x_{21} & x_{22} & \cdots & x_{2k} \\ \vdots & \vdots & & \vdots \\ x_{T1} & x_{T2} & \cdots & x_{Tk} \end{bmatrix} = \begin{bmatrix} X_1' \\ X_2' \\ \vdots \\ X_T' \end{bmatrix},
$$

where X_t is a $k \times 1$ vector:

$$
X_t' = (x_{t1}, x_{t2}, \ldots, x_{tk}), \quad t = 1, \ldots, T.
$$

We wish to represent each observation y_t as a function of $x_{t1}, \ldots, x_{tk}$:

$$
y_t = x_{t1}\beta_1 + x_{t2}\beta_2 + \cdots + x_{tk}\beta_k + \varepsilon_t, \quad t = 1, \ldots, T \tag{1.1}
$$

where ε_t is a "residual" which is left unexplained by the explanatory variables. This model can also be written in the following matrix form:

$$
y = X\beta + \varepsilon \tag{1.2}
$$

where $\varepsilon = (\varepsilon_1, \varepsilon_2, \ldots, \varepsilon_T)'$.

1.2. The least squares problem

1.2.1 In general, we cannot obtain a "perfect fit" $(\varepsilon_t = 0\,, \quad t = 1, \ldots, T)$. In view of this, a natural approach (proposed by Gauss) consists in minimizing the sum of squared residuals:

$$\begin{aligned}\sum_{t=1}^{T} \varepsilon_t^2 &= \sum_{t=1}^{T} [y_t - x_{t1}\beta_1 - \cdots - x_{tk}\beta_k]^2 \\ &= (y - X\beta)'(y - X\beta) \equiv S(\beta)\ .\end{aligned}$$

We consider the problem:

$$\underset{\beta}{\mathrm{Min}}\,(y - X\beta)'(y - X\beta)\ .$$

Since

$$S(\beta) = \left(y' - \beta' X'\right)(y - X\beta) = y'y - 2\beta' X' y + \beta' X' X\beta\,,$$

we have:

$$\frac{\partial S(\beta)}{\partial \beta} = -2X'y + 2X'X\beta\ .$$

To compute the above, we use the following result on differentiation with respect to a vector x :

$$\frac{\partial\,(x'a)}{\partial x} = a\,, \tag{1.3}$$

$$\frac{\partial(x'Ax)}{\partial x} = \left(A + A'\right)x\,. \tag{1.4}$$

For any point $\beta = \hat{\beta}$ such that $S(\beta)$ is a minimum, we must have:

$$\frac{\partial S(\beta)}{\partial \beta}\,|_{\beta=\hat{\beta}} = -2X'y + 2X'X\hat{\beta} = 0$$

hence

$$\left(X'X\right)\hat{\beta} = X'y : \text{normal equations}\,.$$

1.2.2 When $rank(X) = k$, we must have $rank(X'X) = k$ so that $(X'X)^{-1}$ exists. In this case, the normal equations have a unique solution:

$$\hat{\beta} = \left(X'X\right)^{-1} X'y\,. \tag{1.5}$$

Once $\hat{\beta}$ is known, we can compute the "fitted values" and the "residuals" of the model.

1.2.3 The model fitted values are

$$\hat{y} = X\hat{\beta} = X\left(X'X\right)^{-1} X'y = Py\,,$$

where

$$\begin{aligned} P &= X\left(X'X\right)^{-1}X' \qquad \text{(projection matrix)} \\ P' &= P\,,\ PP = P \qquad \text{(symmetric idempotent matrix).} \end{aligned}$$

1.2.4 The model residuals are:

$$\hat{\varepsilon} = y - X\hat{\beta} = y - \hat{y} = y - Py = (I-P)\,y = My$$

where

$$\begin{aligned} PX &= X\,,\ MX = 0\,, && (1.6) \\ PM &= P(I-P) = 0\,,\ MP = 0\,. && (1.7) \end{aligned}$$

1.2.5 Each column of M is orthogonal with each column of X :

$$X'M = 0\,,$$
$$x_i'M = 0\,, \quad i = 1, \ldots, k\,.$$

Residuals and regressors are orthogonal:

$$\begin{aligned} X'\hat{\varepsilon} &= X'My = 0 \\ &\Rightarrow\ x_i'\hat{\varepsilon} = 0\,, \quad i = 1, \ldots, k \\ &\Rightarrow\ i_T'\hat{\varepsilon} = \sum_{t=1}^{T}\hat{\varepsilon}_t = 0\,, \quad \text{if the matrix } X \text{ contains a constant.} \end{aligned}$$

where $\hat{\varepsilon} = (\hat{\varepsilon}_1, \hat{\varepsilon}_2, \ldots, \hat{\varepsilon}_T)'$ et $i_T = (1, 1, \ldots, 1)'$.

1.2.6 Fitted values and residuals are orthogonal:

$$\hat{y}'\hat{\varepsilon} = y'PMy = 0. \tag{1.8}$$

1.2.7 The vector y can be decomposed as the sum of two orthogonal vectors:

$$y = Py + (I-P)\,y = \hat{y} + \hat{\varepsilon}\,. \tag{1.9}$$

1.2.8 For any vector β,

$$\begin{aligned} S(\beta) &\equiv (y - X\beta)'(y - X\beta) = \left(y - X\hat{\beta}\right)'\left(y - X\hat{\beta}\right) + \left(\hat{\beta} - \beta\right)'X'X\left(\hat{\beta} - \beta\right) \\ &\geq \left(y - X\hat{\beta}\right)'\left(y - X\hat{\beta}\right) = S\left(\hat{\beta}\right) \end{aligned}$$

for

$$(y - X\beta)'(y - X\beta) = \left[y - X\hat{\beta} + X\left(\hat{\beta} - \beta\right)\right]'\left[y - X\hat{\beta} + X\left(\hat{\beta} - \beta\right)\right]$$

$$
\begin{aligned}
&= \left[\hat{\varepsilon}+X\left(\hat{\beta}-\beta\right)\right]'\left[\hat{\varepsilon}+X\left(\hat{\beta}-\beta\right)\right] \\
&= \hat{\varepsilon}'\hat{\varepsilon}+2\left(\hat{\beta}-\beta\right)'X'\hat{\varepsilon}+\left(\hat{\beta}-\beta\right)'X'X\left(\hat{\beta}-\beta\right) \\
&= \hat{\varepsilon}'\hat{\varepsilon}+\left(\hat{\beta}-\beta\right)'X'X\left(\hat{\beta}-\beta\right)\ .
\end{aligned}
$$

This directly verifies that $\beta=\hat{\beta}$ minimizes $S(\beta)$.

2. Classical linear model

In order to establish the statistical properties of $\hat{\beta}$, we need assumptions on X and ε. The following assumptions define the *classical linear model* (CLM).

2.1 Assumption $y=X\beta+\varepsilon$
where y is a $T\times 1$ vector of observations on a dependent variable ,
X is a $T\times k$ matrix of observations on explanatory variables,
β is a $k\times 1$ vector of fixed parameters,
ε is a $T\times 1$ vector of random disturbances.

2.2 Assumption $E(\varepsilon)=0$.

2.3 Assumption $E[\varepsilon\varepsilon']=\sigma^2 I_T$.

2.4 Assumption *X is fixed (non-stochastic).*

2.5 Assumption $\operatorname{rank}(X)=k<T$.

From the assumption 2.1 - 2.4, we see that:

$$
\begin{aligned}
\mathsf{E}(y) &= \mathsf{E}(y\mid X)=X\beta=\begin{pmatrix} X_1'\beta \\ \vdots \\ X_T'\beta \end{pmatrix} \\
&= (x_1,x_2,\ldots,x_k)\begin{pmatrix} \beta_1 \\ \beta_2 \\ \vdots \\ \beta_k \end{pmatrix} \\
&= x_1\beta_1+x_2\beta_2+\cdots+x_k\beta_k\,, \\
\mathsf{V}(y) &= \mathsf{V}(y\mid X)=\sigma^2 I_T \\
&= \begin{bmatrix} \sigma^2 & 0 & \cdots & 0 \\ 0 & \sigma^2 & \cdots & 0 \\ \vdots & \vdots & & \vdots \\ 0 & 0 & \cdots & \sigma^2 \end{bmatrix}=\mathsf{V}(\varepsilon)\ .
\end{aligned}
$$

If, furthermore, we add the assumption that ε follows a multinormal (or Gaussian) distribution, we get the normal classical linear model (NCLM).

2.6 Assumption *ε follows a multinormal distribution.*

3. Linear unbiased estimation

From the assumptions 2.1 - 2.5, we can make the following observations.

3.1 $\hat{\beta}$ is linear with respect to y.

PROOF $\hat{\beta}$ has the form $\hat{\beta} = Ay$, where $A = (X'X)^{-1}X'$ is a non-stochastic matrix. □

3.2 $\hat{\beta} = (X'X)^{-1}X'(X\beta + \varepsilon) = \beta + (X'X)^{-1}X'\varepsilon$.

3.3 $\hat{\beta}$ is an unbiased estimator of β.

PROOF $E(\hat{\beta}) = \beta + (X'X)^{-1}X'E(\varepsilon) = \beta$. □

3.4 $\mathsf{V}(\hat{\beta}) = \sigma^2(X'X)^{-1}$.

PROOF

$$\begin{aligned}\mathsf{V}(\hat{\beta}) &= \mathsf{E}\left[(\hat{\beta}-\beta)(\hat{\beta}-\beta)'\right] \\ &= \mathsf{E}\left[\left(X'X\right)^{-1}X'\varepsilon\varepsilon'X\left(X'X\right)^{-1}\right] \\ &= \left(X'X\right)^{-1}X'\mathsf{E}\left(\varepsilon\varepsilon'\right)X\left(X'X\right)^{-1} \\ &= \sigma^2\left(X'X\right)^{-1}\end{aligned}$$

where the last identity follows from Assumption 2.3. □

3.5 Theorem GAUSS-MARKOV THEOREM. *$\hat{\beta}$ is the best estimator of β in the class of linear linear unbiased estimators $(BLUE)$ of β, i.e. $\mathsf{V}(\tilde{\beta}) - \mathsf{V}(\hat{\beta})$ is a positive semidefinite matrix for any linear unbiased estimator (LUE) $\tilde{\beta}$ of β. In particular, if $\tilde{\beta} = Cy$ and $D = C - (X'X)^{-1}X'$, then*

$$\mathsf{V}(\tilde{\beta}) = \mathsf{V}(\hat{\beta}) + \sigma^2 DD'.$$

PROOF Since $\tilde{\beta}$ is unbiased and

$$C = D + \left(X'X\right)^{-1} X',$$

we have:

$$\begin{aligned}
\mathsf{E}(\tilde{\beta}) &= \mathsf{E}\left\{\left[D + \left(X'X\right)^{-1} X'\right](X\beta + \varepsilon)\right\} \\
&= DX\beta + \beta \\
&= \beta,
\end{aligned}$$

hence

$$DX = 0 \quad \text{and} \quad CX = I_k .$$

Consequently,

$$\tilde{\beta} = Cy = CX\beta + C\varepsilon = \beta + C\varepsilon$$

and

$$\tilde{\beta} - \beta = C\varepsilon,$$

hence

$$\begin{aligned}
\mathsf{V}(\tilde{\beta}) &= \mathsf{E}\left[(\tilde{\beta} - \beta)(\tilde{\beta} - \beta)'\right] = \mathsf{E}\left[C\varepsilon\varepsilon'C'\right] = \sigma^2 CC' \\
&= \sigma^2\left[D + \left(X'X\right)^{-1} X'\right]\left[D' + X\left(X'X\right)^{-1}\right] \\
&= \sigma^2\left[DD' + \left(X'X\right)^{-1}\right] = \sigma^2 DD' + \sigma^2\left(X'X\right)^{-1} \\
&= \sigma^2 DD' + \mathsf{V}(\hat{\beta})
\end{aligned}$$

and

$$\mathsf{V}(\tilde{\beta}) - \mathsf{V}(\hat{\beta}) = \sigma^2 DD' \tag{3.1}$$

is a positive semidefinite matrix. □

3.6 Corollary *Let w be a $k \times 1$ vector of constants. Then,*

$$\mathsf{V}(w'\tilde{\beta}) \geq \mathsf{V}(w'\hat{\beta})$$

for any linear unbiased estimator $\tilde{\beta}$ of β.

PROOF Since $E\left(\tilde{\beta}\right) = E\left(\hat{\beta}\right) = \beta$, we have:

$$\begin{aligned}
\mathsf{E}\left(w'\tilde{\beta}\right) &= \mathsf{E}\left(w'\hat{\beta}\right) = w'\beta, \\
\mathsf{V}\left(w'\tilde{\beta}\right) &= w'\mathsf{V}\left(\tilde{\beta}\right)w = w'\left[\sigma^2 DD' + \mathsf{V}\left(\hat{\beta}\right)\right]w \\
&= \sigma^2 w'DD'w + w'\mathsf{V}\left(\hat{\beta}\right)w
\end{aligned}$$

$$= \sigma^2 w'DD'w + \mathsf{V}\left(w'\hat{\beta}\right) \geq \mathsf{V}\left(w'\hat{\beta}\right) ,$$

for $w'DD'w \geq 0$. $\square$

In particular, we must have:

$$\mathsf{V}(\tilde{\beta}_i) \geq \mathsf{V}(\hat{\beta}_i) , \quad i = 1, \ldots k .$$

3.7 Theorem GENERALIZED GAUSS-MARKOV THEOREM. *Let L be a $r \times k$ fixed matrix and $\gamma = L\beta$. Then $\hat{\gamma} = L\hat{\beta}$ is the BLUE γ, i.e. $\mathsf{V}(\tilde{\gamma}) - \mathsf{V}(\hat{\gamma})$ is a positive semidefinite matrix for any linear unbiased estimator $\tilde{\gamma}$ of γ. In particular, if $\tilde{\gamma} = Cy$ and $D = C - L(X'X)^{-1}X'$, then*

$$\mathsf{V}(\tilde{\gamma}) = \mathsf{V}(\hat{\gamma}) + \sigma^2 DD'$$

and

$$\mathsf{C}(\tilde{\gamma} - \hat{\gamma}, \hat{\gamma}) = 0 .$$

PROOF Since $\tilde{\gamma}$ is unbiased and

$$C = D + L\left(X'X\right)^{-1}X'$$

we have

$$\begin{aligned} \mathsf{E}(\tilde{\gamma}) &= \mathsf{E}\left\{\left(D + L\left(X'X\right)^{-1}X'\right](X\beta + \varepsilon)\right\} \\ &= DX\beta + L\beta = DX\beta + \gamma \\ &= \gamma , \end{aligned}$$

hence

$$DX = 0 \quad \text{and} \quad CX = L .$$

Consequently,

$$\begin{aligned} \tilde{\gamma} &= Cy = CX\beta + C\varepsilon \\ &= L\beta + C\varepsilon = \gamma + C\varepsilon \end{aligned}$$

and

$$\begin{aligned} \mathsf{V}(\tilde{\gamma}) &= \mathsf{E}\left[(\tilde{\gamma} - \gamma)(\tilde{\gamma} - \gamma)'\right] = \mathsf{E}\left[C\varepsilon\varepsilon'C'\right] = \sigma^2 CC' \\ &= \sigma^2\left[D + L\left(X'X\right)^{-1}X'\right]\left[D' + X\left(X'X\right)^{-1}L'\right] \\ &= \sigma^2\left[DD' + L\left(X'X\right)^{-1}L'\right] \\ &= \sigma^2 DD' + \sigma^2 L\left(X'X\right)^{-1}L' = \sigma^2 DD' + \mathsf{V}(L\hat{\beta}) \\ &= \sigma^2 DD' + \mathsf{V}(\hat{\gamma}) , \end{aligned}$$

so

$$\mathsf{V}(\tilde{\gamma}) - \mathsf{V}(\hat{\gamma}) = \sigma^2 DD' \tag{3.2}$$

is a positive semidefinite matrix, and

$$\begin{aligned}
\mathsf{C}(\tilde{\gamma}, \hat{\gamma}) &= \mathsf{E}\left[C\varepsilon\varepsilon' X \left(X'X\right)^{-1} L'\right] \\
&= \sigma^2 CX \left(X'X\right)^{-1} L' = \sigma^2 L \left(X'X\right)^{-1} L' = \mathsf{V}(\hat{\gamma}) \,,
\end{aligned}$$

$$\mathsf{C}(\tilde{\gamma} - \hat{\gamma}, \hat{\gamma}) = \mathsf{C}(\tilde{\gamma}, \hat{\gamma}) - \mathsf{C}(\hat{\gamma}, \hat{\gamma}) = \mathsf{V}(\hat{\gamma}) - \mathsf{V}(\hat{\gamma}) = 0 \,. \tag{3.3}$$

□

3.8 Corollary QUADRATIC GAUSS-MARKOV OPTIMALITY. *Let Q be a $r \times r$ positive semidefinite fixed matrix and L a $r \times k$ fixed matrix, $\gamma = L\beta$ and $\hat{\gamma} = L\hat{\beta}$. Then*

$$\mathsf{E}\left[(\tilde{\gamma} - \gamma)' Q (\tilde{\gamma} - \gamma)\right] \geq \mathsf{E}\left[(\hat{\gamma} - \gamma)' Q (\hat{\gamma} - \gamma)\right]$$

for any linear unbiased estimator $\tilde{\gamma}$ of γ.

PROOF Let $\tilde{\gamma} = C\gamma$ and $D = C - L\,(X'X)^{-1} X'$. Then

$$\begin{aligned}
\mathsf{E}\left[(\tilde{\gamma} - \gamma)' Q (\tilde{\gamma} - \gamma)\right] &= \mathsf{E}\left[\operatorname{tr} Q\,(\tilde{\gamma} - \gamma)\,(\tilde{\gamma} - \gamma)'\right] \\
&= \operatorname{tr} Q \mathsf{E}\left[(\tilde{\gamma} - \gamma)\,(\tilde{\gamma} - \gamma)'\right] \\
&= \operatorname{tr} Q\left[\sigma^2 DD' + \mathsf{V}(\hat{\gamma})\right] \\
&= \sigma^2 \operatorname{tr}\left(QDD'\right) + \operatorname{tr}\left[Q\mathsf{V}(\hat{\gamma})\right] \\
&= \sigma^2 \operatorname{tr}\left(D'QD\right) + \operatorname{tr} Q \mathsf{E}\left[(\hat{\gamma} - \gamma)\,(\hat{\gamma} - \gamma)'\right] \\
&= \sigma^2 \operatorname{tr}\left(D'QD\right) + \mathsf{E}\left[\operatorname{tr}\,(\hat{\gamma} - \gamma)' Q\,(\hat{\gamma} - \gamma)\right] \\
&= \sigma^2 \operatorname{tr}\left(D'QD\right) + \mathsf{E}\left[(\hat{\gamma} - \gamma)' Q\,(\hat{\gamma} - \gamma)\right] \\
&\geq \mathsf{E}\left[(\hat{\gamma} - \gamma)' Q\,(\hat{\gamma} - \gamma)\right]
\end{aligned}$$

since Q is p.s.d. $\Rightarrow D'QD$ is p.s.d. $\Rightarrow \operatorname{tr} D'QD \geq 0$. □

3.9 Corollary *For any LUE of $\tilde{\gamma}$ of $\gamma = L\beta$,*

$$\operatorname{tr} \mathsf{V}(\tilde{\gamma}) \geq \operatorname{tr} \mathsf{V}(\hat{\gamma}) \;.$$

PROOF

$$\operatorname{tr} \mathsf{V}(\tilde{\gamma}) = \operatorname{tr} \mathsf{E}\left[(\tilde{\gamma} - \gamma)\,(\tilde{\gamma} - \gamma)'\right] = \mathsf{E}\left[\operatorname{tr}\,(\tilde{\gamma} - \gamma)\,(\tilde{\gamma} - \gamma)'\right]$$

$$= \mathsf{E}\left[(\tilde{\gamma}-\gamma)'(\tilde{\gamma}-\gamma)\right] \geq \mathsf{E}\left[(\hat{\gamma}-\gamma)'(\hat{\gamma}-\gamma)\right] = \operatorname{tr}\mathsf{V}(\hat{\gamma})$$

by Corollary 3.8 with $Q = I$. □

3.10 Lemma PROPERTIES OF MATRIX DOMINANCE. *If $A = B + C$ where B is a p.d. matrix and C is a p.s.d. matrix, then*

(a) *A is p.d.,*

(b) $|B| \leq |A|$,

(c) $B^{-1} - A^{-1}$ *is p.s.d.*

3.11 Corollary *Let L be an $r \times k$ fixed matrix, $\gamma = L\beta$ and $\hat{\gamma} = L\hat{\beta}$. Then*

$$|\mathsf{V}(\tilde{\gamma})| \geq |\mathsf{V}(\hat{\gamma})|$$

for any LUE $\tilde{\gamma}$ of γ.

PROOF Since $\hat{\gamma}$ is the BLUE of γ (by the generalized Gauss-Markov theorem), we have:

$$\mathsf{V}(\tilde{\gamma}) = \mathsf{V}(\hat{\gamma}) + C \tag{3.4}$$

where C is p.s.d. If $|\mathsf{V}(\hat{\gamma})| = 0$, then $|\mathsf{V}(\hat{\gamma})| \leq |\mathsf{V}(\tilde{\gamma})|$, for car $|\mathsf{V}(\tilde{\gamma})| \geq 0$. If $|\mathsf{V}(\hat{\gamma})| > 0$, then $\mathsf{V}(\hat{\gamma})$ is p.d. This entails that $\mathsf{V}(\tilde{\gamma})$ is also p.d. and $|\mathsf{V}(\hat{\gamma})| \leq |\mathsf{V}(\tilde{\gamma})|$. □

3.12 $\hat{y} = X\beta + P\varepsilon$, $\hat{\varepsilon} = My = M\varepsilon$.

PROOF

$$\begin{aligned}
\hat{y} &= Py = P[X\beta + \varepsilon] = X\beta + P\varepsilon\,, \quad \text{car } PX = X\,, \\
\hat{\varepsilon} &= My = M[X\beta + \varepsilon] = M\varepsilon\,, \quad \text{car } MX = 0\,.
\end{aligned}$$

□

3.13 $E(\hat{y}) = X\beta$, $E(\hat{\varepsilon}) = 0$.

PROOF

$$\begin{aligned}
\mathsf{E}(\hat{\gamma}) &= \mathsf{E}[X\beta + P\varepsilon] = X\beta + P\mathsf{E}(\varepsilon) = X\beta\,, \\
\mathsf{E}(\hat{\varepsilon}) &= \mathsf{E}(y - \hat{y}) = X\beta - X\beta = 0\,.
\end{aligned}$$

□

3.14 $\mathsf{V}(\hat{y}) = \sigma^2 P$, $\mathsf{V}(\hat{\varepsilon}) = \sigma^2 M$.

PROOF

$$\begin{aligned} \mathsf{V}(\hat{y}) &= \mathsf{V}\left(X\hat{\beta}\right) = X\mathsf{V}\left(\hat{\beta}\right)X' = \sigma^2 X\left(X'X\right)^{-1}X' = \sigma^2 P \ , \\ \mathsf{V}(\hat{\varepsilon}) &= \mathsf{V}(My) = M\mathsf{V}(y)M' = \sigma^2 M \ . \end{aligned}$$

□

3.15 $\hat{y}$ is the best linear unbiased estimator of $X\beta$.

PROOF This follows directly on taking $L = X$ in the generalized Gauss-Markov theorem. □

3.16 $\hat{\varepsilon}$ is the best linear unbiased estimator (BLUE) of ε, in the sense that $E(\hat{\varepsilon} - \varepsilon) = 0$ and

$$\mathsf{V}(\tilde{\varepsilon} - \varepsilon) - \mathsf{V}(\hat{\varepsilon} - \varepsilon) \text{ is a p.s.d. matrix}$$

for for LUE $\tilde{\varepsilon}$ of ε.

PROOF Since $\tilde{\varepsilon}$ is a LUE of ε, we must have:

$$\tilde{\varepsilon} = Ay \quad \text{and} \quad \mathsf{E}(\tilde{\varepsilon} - \varepsilon) = 0 \ .$$

Consequently,

$$\begin{aligned} \mathsf{E}(\tilde{\varepsilon}) &= \mathsf{E}(Ay) \\ &= \mathsf{E}[A(X\beta + \varepsilon)] = AX\beta = 0 \ , \forall \beta \ , \end{aligned}$$

which entails that

$$AX = 0 \,,$$
$$\tilde{\varepsilon} = A(X\beta + \varepsilon) = A\varepsilon \ .$$

Let

$$B = A - M \quad \text{where} \quad M = I - X\left(X'X\right)^{-1}X' \ .$$

Then

$$AX = [B + M]X = BX = 0 \ , \quad \text{since} \quad MX = 0 \ ,$$

hence

$$\mathsf{V}(\tilde{\varepsilon} - \varepsilon) = \mathsf{V}[A\varepsilon - \varepsilon]$$

$$
\begin{aligned}
&= \mathsf{V}\left[(B+M)\varepsilon-\varepsilon\right] = \mathsf{V}\left[(B+M-I)\varepsilon\right] \\
&= \mathsf{E}\left[(B+M-I)\varepsilon\varepsilon'\left(B'+M-I\right)\right] \\
&= \sigma^2\left[B-X\left(X'X\right)^{-1}X'\right]\left[B'-X\left(X'X\right)^{-1}X'\right] \\
&= \sigma^2\left[BB'+X\left(X'X\right)^{-1}X'\right] ,
\end{aligned}
$$

and

$$
\begin{aligned}
\mathsf{V}(\hat{\varepsilon}-\varepsilon) &= \mathsf{E}\left[(M-I)\varepsilon\varepsilon'(M-I)\right] \\
&= \sigma^2(I-M) = \sigma^2 X\left(X'X\right)^{-1}X' ,
\end{aligned}
$$

so that

$$\mathsf{V}(\tilde{\varepsilon}-\varepsilon) = \sigma^2 BB' + \mathsf{V}(\hat{\varepsilon}-\varepsilon) \ .$$

Thus

$$\mathsf{V}(\tilde{\varepsilon}-\varepsilon) - \mathsf{V}(\hat{\varepsilon}-\varepsilon) = \sigma^2 BB'$$

a p.s.d. matrix. □

3.17 $\mathsf{C}(\hat{\beta}, \hat{\varepsilon}) = \mathsf{C}(\hat{\beta}, y-X\hat{\beta}) = 0$.

PROOF

$$
\begin{aligned}
\mathsf{C}(\hat{\beta}, \hat{\varepsilon}) &= \mathsf{E}\left[(\hat{\beta}-\beta)\hat{\varepsilon}'\right] = \mathsf{E}\left[\left(X'X\right)^{-1}X'\varepsilon\varepsilon'M\right] \\
&= \sigma^2\left(X'X\right)^{-1}X'M = 0 \ .
\end{aligned}
$$

□

3.18 $\mathsf{C}(\hat{y}, \hat{\varepsilon}) = 0$.

PROOF

$$
\begin{aligned}
\mathsf{C}(\hat{y}, \hat{\varepsilon}) &= \mathsf{E}\left[(X\hat{\beta}-X\beta)\hat{\varepsilon}'\right] \\
&= X\,\mathsf{E}\left[(\hat{\beta}-\beta)\hat{\varepsilon}'\right] = X\,\mathsf{C}(\hat{\beta}, \hat{\varepsilon}) = 0 \ .
\end{aligned}
$$

□

3.19 Estimation of σ^2. Since $\sigma^2 = E\left(\varepsilon_t^2\right), t = 1, \ldots, T$, it is natural to consider the residuals of the regression which can be viewed as estimations of the error terms ε_t:

$$\hat{\varepsilon} = y - X\hat{\beta} = My = M(X\beta+\varepsilon) = M\varepsilon \ ,$$

$$\sum_{t=1}^{T}\hat{\varepsilon}_t^2=\hat{\varepsilon}'\hat{\varepsilon}=\varepsilon'M'M\varepsilon=\varepsilon'M\varepsilon\ ,$$

hence

$$\begin{aligned}\mathsf{E}\left[\hat{\varepsilon}'\hat{\varepsilon}\right] &= \mathsf{E}\left[\varepsilon'M\varepsilon\right]=\mathsf{E}\left[\mathrm{tr}\left(\varepsilon'M\varepsilon\right)\right]\\ &= \mathsf{E}\left[\mathrm{tr}\left(M\varepsilon\varepsilon'\right)\right]=\mathrm{tr}\left[M\mathsf{E}\left(\varepsilon\varepsilon'\right)\right]\\ &= \sigma^2\mathrm{tr}M\ ,\end{aligned}$$

where

$$\begin{aligned}\mathrm{tr}M &= \mathrm{tr}\left[I_T-X\left(X'X\right)^{-1}X'\right]=\mathrm{tr}I_T-\mathrm{tr}\left[X\left(X'X\right)^{-1}X'\right]\\ &= \mathrm{tr}I_T-\mathrm{tr}\left[X'X\left(X'X\right)^{-1}\right]=\mathrm{tr}I_T-\mathrm{tr}I_k\\ &= T-k\ .\end{aligned}$$

Thus,

$$\begin{aligned}\mathsf{E}\left(\hat{\varepsilon}'\hat{\varepsilon}\right) &= \sigma^2\left(T-k\right)\\ \mathsf{E}\left[\frac{\hat{\varepsilon}'\hat{\varepsilon}}{T-k}\right] &= \sigma^2\ .\end{aligned}$$

3.20 The statistic

$$s^2=\hat{\varepsilon}'\hat{\varepsilon}/\left(T-k\right)=y'My/\left(T-k\right)$$

is an unbiased estimator of σ^2, and $s^2\left(X'X\right)^{-1}$ is an unbiased estimator of $\mathsf{V}\left(\hat{\beta}\right)=\sigma^2\left(X'X\right)^{-1}$:

$$\begin{aligned}\mathsf{E}\left(s^2\right) &= \sigma^2\ ,\\ \mathsf{E}\left[s^2\left(X'X\right)^{-1}\right] &= \sigma^2\left(X'X\right)^{-1}\ .\end{aligned}$$

4. Prediction

In the previous section, we studied how one can estimate β in the linear regression model. Suppose now we know the matrix X_0 of explanatory variables for m additional periods (or observations). We wish to predict the corresponding values of y:

$$y_0=X_0\beta+\varepsilon_0$$

where

$$\mathsf{E}\left(\varepsilon_0\right)=0\ ,\mathsf{V}\left(\varepsilon_0\right)=\sigma^2I_m\ ,\mathsf{E}\left(\varepsilon\varepsilon_0'\right)=0\ .$$

The natural "predictor" in this case is:

$$\hat{y}_0=X_0\hat{\beta}=X_0\left(X'X\right)^{-1}X'y\ . \tag{4.1}$$

We can then show the following properties.

4.1 $\hat{y}_0$ is an unbiased estimator of $X_0\beta$:

$$\mathsf{E}(\hat{y}_0) = X_0\beta = \mathsf{E}(y_0), \quad \mathsf{E}(\hat{y}_0 - y_0) = 0.$$

4.2 $\mathsf{V}(\hat{y}_0) = \mathsf{V}\left(X_0\hat{\beta}\right) = X_0\mathsf{V}\left(\hat{\beta}\right)X_0' = \sigma^2 X_0 (X'X)^{-1} X_0'.$

4.3 $\mathsf{C}(y_0, \hat{y}_0) = 0.$

PROOF

$$\begin{aligned}
\mathsf{C}(y_0, \hat{y}_0) &= \mathsf{E}\left[(y_0 - X_0\beta)\left(X_0\hat{\beta} - X_0\beta\right)'\right] \\
&= \mathsf{E}\left[\varepsilon_0\left(\hat{\beta} - \beta\right)' X_0'\right] = \mathsf{E}\left[\varepsilon_0\varepsilon' X\left(X'X\right)^{-1} X_0'\right] = 0 .
\end{aligned}$$

□

4.4 $\hat{y}_0$ is best linear unbiased estimator of $X_0\beta$, in the sense that $\mathsf{V}(\tilde{y}_0) - \mathsf{V}(\hat{y}_0)$ is a p.s.d. matrix for any linear unbiased estimator $\tilde{y}_0$ of $X_0\beta$. In particular, if $\tilde{y}_0 = Cy$ and $D = C - X_0(X'X)^{-1}X'$, then

$$\mathsf{V}(\tilde{y}_0) = \mathsf{V}(\hat{y}_0) + \sigma^2 DD' .$$

PROOF This follows directly from the generalized Gauss-Markov theorem. □

The "prediction errors" are given by:

$$\begin{aligned}
\hat{e}_0 &= y_0 - \hat{y}_0 = y_0 - X_0\hat{\beta} \\
&= X_0\beta + \varepsilon_0 - X_0\hat{\beta} = \varepsilon_0 + X_0\left(\beta - \hat{\beta}\right) .
\end{aligned}$$

4.5 $\hat{y}_0$ is a linear unbiased predictor (LUP) of y_0:

$$\mathsf{E}[\hat{e}_0] = 0 .$$

PROOF $\hat{y}_0 = X_0\hat{\beta}$ and

$$\mathsf{E}[\hat{e}_0] = \mathsf{E}[y_0 - \hat{y}_0] = X_0\beta - X_0\beta = 0 .$$

□

4.6 $\mathsf{V}(\hat{e}_0) = \sigma^2 \left[I_m + X_0 (X'X)^{-1} X_0' \right]$.

PROOF

$$\begin{aligned}\mathsf{V}(y_0 - \hat{y}_0) &= \mathsf{V}(y_0) + \mathsf{V}(\hat{y}_0) - \mathsf{C}(y_0, \hat{y}_0) - \mathsf{C}(\hat{y}_0, y_0) \\ &= \sigma^2 I_m + \sigma^2 X_0 \left(X'X\right)^{-1} X_0' \\ &= \sigma^2 \left[I_m + X_0 \left(X'X\right)^{-1} X_0' \right] .\end{aligned}$$

□

4.7 Theorem *$\hat{y}_0$ is the best linear unbiased predictor (BLUP) of y_0, in the sense that $\mathsf{V}(y_0 - \tilde{y}_0) - \mathsf{V}(y_0 - \hat{y}_0)$ is a p.s.d. matrix for any LUP $\tilde{y}_0$ of y_0. In particular, if $\tilde{y}_0 = Cy$ and $D = C - X_0 (X'X)^{-1} X'$, then*

$$\mathsf{V}(y_0 - \tilde{y}_0) = \mathsf{V}(y_0 - \hat{y}_0) + \sigma^2 DD' .$$

PROOF

$$\mathsf{V}(y_0 - \tilde{y}_0) = \mathsf{V}(y_0) + \mathsf{V}(\tilde{y}_0) - \mathsf{C}(y_0, \tilde{y}_0) - \mathsf{C}(\tilde{y}_0, y_0)$$

where

$$\mathsf{C}(y_0, \tilde{y}_0) = \mathsf{E}\left[\varepsilon_0 \varepsilon' C'\right] = 0$$

for, by the generalized Gauss-Markov theorem,

$$\mathsf{E}[\tilde{y}_0] = X_0\beta \Rightarrow CX = X_0 \Rightarrow \tilde{y}_0 = C(X\beta + \varepsilon) = X_0\beta + C\varepsilon .$$

Further, $\mathsf{V}(\tilde{y}_0) = \mathsf{V}(\hat{y}_0) + \sigma^2 DD'$ and $\mathsf{V}(y_0) = \sigma^2 I_m$. Consequently,

$$\begin{aligned}\mathsf{V}(y_0 - \tilde{y}_0) &= \sigma^2 I_m + \mathsf{V}(\hat{y}_0) + \sigma^2 DD' \\ &= \left[\sigma^2 I_m + \sigma^2 X_0 \left(X'X\right)^{-1} X_0'\right] + \sigma^2 DD' \\ &= \mathsf{V}(y_0 - \hat{y}_0) + \sigma^2 DD' .\end{aligned}$$

□

5. Estimation with Gaussian errors

If we wish to build confidence intervals and perform hypothesis tests, we need a more complete specification of the error distribution. The standard hypothesis for this is to assume that the errors follow a Gaussian distribution.

5.1 Assumption $\varepsilon \sim N_T\left[0, \sigma^2 I_T\right]$.

This means that the errors ε_t are i.i.d. $N\left[0, \sigma^2\right]$. We can now completely establish the distribution of the least squares estimator.

5.2 $y \sim N\left[X\beta, \sigma^2 I_T\right]$, since $y = X\beta + \varepsilon$.

5.3 $\hat{\beta} \sim N\left[\beta, \sigma^2 \left(X'X\right)^{-1}\right]$, since $\hat{\beta} = \left(X'X\right)^{-1} X'y$.

The probability density function of y is given by:

$$L\left(y; X\beta, \sigma^2 I_T\right) = \frac{1}{\left(2\pi\sigma^2\right)^{T/2}} \exp\left\{-\frac{1}{2}\frac{(y - X\beta)'(y - X\beta)}{\sigma^2}\right\} .$$

5.4 $\hat{\beta} = \left(X'X\right)^{-1} X'y$ and $\hat{\sigma}^2 = \hat{\varepsilon}'\hat{\varepsilon}/T$ are the maximum likelihood estimators of β and σ^2 respectively.

PROOF To maximize L is equivalent to maximizing $\ln(L)$. Since

$$\begin{aligned}
\ln(L) &= -\frac{T}{2}\ln(2\pi) - \frac{T}{2}\ln(\sigma^2) - \frac{1}{2\sigma^2}(y - X\beta)'(y - X\beta) \\
&= -\frac{T}{2}\ln(2\pi) - \frac{T}{2}\ln(\sigma^2) - \frac{1}{2\sigma^2}\left[y'y - 2y'X\beta + \beta'X'X\beta\right] ,
\end{aligned}$$

the first-order conditions (which are necessary) for a maximum is:

$$\begin{aligned}
\frac{\partial\left(\ln(L)\right)}{\partial\beta} &= -\frac{1}{2\sigma^2}\left[-2X'y + 2\left(X'X\right)\beta\right] = 0 , \\
\frac{\partial\left(\ln(L)\right)}{\partial\sigma^2} &= -\frac{T}{2}\frac{1}{\sigma^2} + \frac{1}{2\sigma^4}(y - X\beta)'(y - X\beta) = 0 ,
\end{aligned}$$

hence

$$\begin{aligned}
\left(X'X\right)\hat{\beta} &= X'y \, , \hat{\beta} = \left(X'X\right)^{-1} X'y , \\
\hat{\sigma}^2 &= \left(y - X\hat{\beta}\right)'\left(y - X\hat{\beta}\right)/T .
\end{aligned}$$

Further the second-order derivative of $\ln(L)$ is:

$$\frac{\partial\left(\ln(L)\right)}{\partial\beta'\partial\beta} = -\frac{1}{\sigma^2}\left(X'X\right) \tag{5.1}$$

which is negative semidefinite as required for a maximum. □

5.5 $\hat{y} = X\hat{\beta} \sim N_T\left[X\beta, \sigma^2 P\right]$.

5.6 $\hat{\varepsilon} = M\varepsilon \sim N_T\left[0, \sigma^2 M\right]$.

5.7 $\hat{\varepsilon}$ and $\hat{\beta}$ are independent, because $\hat{\varepsilon}$ et $\hat{\beta}$ are multinormal and $\mathsf{C}(\hat{\beta}, \hat{\varepsilon}) = 0$.

5.8 $\hat{\varepsilon}$ and $\hat{y}$ are independent, because $\hat{\varepsilon}$ and $\hat{y}$ are multinormal and $\mathsf{C}(\hat{y}, \hat{\varepsilon}) = 0$.

5.9 Lemma DISTRIBUTION OF AN IDEMPOTENT QUADRATIC FORM IN I.I.D. GAUSSIAN VARIABLES. *Let Q be a $T \times T$ symmetric idempotent matrix of rank $q \leq T$. If $\varepsilon \sim N_T\left[0, \sigma^2 I_T\right]$, then*

$$\varepsilon' Q \varepsilon / \sigma^2 \sim \chi^2(q) \ .$$

PROOF Since Q is a symmetric idempotent matrix, there is a $T \times T$ orthogonal matrix C, i.e. $CC' = C'C = I_T$, such that

$$CQC' = \begin{pmatrix} I_q & 0 \\ 0 & 0 \end{pmatrix} ,$$

hence

$$\varepsilon' Q \varepsilon = \varepsilon' C' C Q C' C \varepsilon = (C\varepsilon)' \left(CQC'\right)(C\varepsilon) \ .$$

Further,

$$\begin{aligned} \varepsilon \quad &\sim \quad N\left[0, \sigma^2 I_T\right] \Rightarrow C\varepsilon \sim N\left[0, \sigma^2 C I_T C'\right] \\ &\Rightarrow \quad C\varepsilon \sim N\left[0, \sigma^2 I_T\right] \ . \end{aligned}$$

Let $v = C\varepsilon = (v_1, v_2, \ldots, v_T)'$. Then

$$v_1, v_2, \ldots, v_T \text{ are i.i.d. } N\left[0, \sigma^2\right]$$

and

$$\begin{aligned} \varepsilon' Q \varepsilon &= v'\left(CQC'\right) v \\ &= (v_1, v_2, \ldots, v_T) \begin{pmatrix} I_q & 0 \\ 0 & 0 \end{pmatrix} \begin{bmatrix} v_1 \\ v_2 \\ \vdots \\ v_T \end{bmatrix} \\ &= v_1^2 + v_2^2 + \cdots + v_q^2 + 0\ .v_{q+1}^2 \cdots + 0\ .v_T^2 \\ &= \sum_{t=1}^{q} v_t^2 \ . \end{aligned}$$

This entails

$$\frac{\varepsilon' Q \varepsilon}{\sigma^2} = \sum_{t=1}^{q} \left(\frac{v_t}{\sigma}\right)^2 ,$$

$$\text{where } \frac{v_t}{\sigma} \overset{ind}{\sim} N[0,1] \ , \quad t = 1, \ldots, T \ ,$$

and

$$\varepsilon' Q \varepsilon / \sigma^2 \sim \chi^2(q) \ .$$

□

5.10

$$\frac{S\left(\hat{\beta}\right)}{\sigma^2} = \frac{\hat{\varepsilon}'\hat{\varepsilon}}{\sigma^2} \sim \chi^2(T-k) \ .$$

PROOF This follows directly on applying Lemma 5.9 with $Q = M$ and the fact that $\operatorname{tr}(M) = T - k$. □

5.11 Let R be a $q \times k$ fixed matrix. Then,

$$R\hat{\beta} \sim N_q\left[R\beta, \sigma^2 R\left(X'X\right)^{-1} R'\right] \ . \tag{5.2}$$

Further $R\hat{\beta}$ and s^2 are independent.

PROOF $\hat{\beta} \sim N\left[\beta, \sigma^2 (X'X)^{-1}\right]$ entails $R\hat{\beta} \sim N\left[R\beta, \sigma^2 R(X'X)^{-1} R'\right]$. Since $\hat{\beta}$ and $\hat{\varepsilon}$ are independent, $R\hat{\beta}$ and $\hat{\varepsilon}'\hat{\varepsilon}$ are also independent, so that $R\hat{\beta}$ and $s^2 = \hat{\varepsilon}'\hat{\varepsilon}/(T-k)$ are independent. □

5.12 Let R be a $q \times k$ fixed matrix of rank q, $r = R\beta$ and

$$S(R, \hat{\beta}) = [R\hat{\beta} - r]' \left[R\left(X'X\right)^{-1} R'\right]^{-1} [R\hat{\beta} - r] \ .$$

Then

$$S(R, \hat{\beta})/\sigma^2 \sim \chi^2(q) \ . \tag{5.3}$$

Further, $S(R, \hat{\beta})$ and s^2 are independent.

PROOF

$$R\hat{\beta} - r = R\left(\hat{\beta} - \beta\right)$$

and

$$R\left(\hat{\beta} - \beta\right) \sim N_q\left[0, \sigma^2 R\left(X'X\right)^{-1} R'\right] \ .$$

Thus,

$$\begin{aligned} S(R,\hat{\beta})/\sigma^2 &= \left[R\left(\hat{\beta}-\beta\right)\right]'\left[\sigma^2 R\left(X'X\right)^{-1}R'\right]^{-1}\left[R\left(\hat{\beta}-\beta\right)\right] \\ &\sim \chi^2(q)\ . \end{aligned}$$

$\square$

6. Confidence and prediction intervals

6.1. Confidence interval for the error variance

In the normal classical linear model, we have:

$$\hat{\varepsilon}'\hat{\varepsilon}/\sigma^2 = (T-k)s^2/\sigma^2 \sim \chi^2(T-k)\ .$$

Thus, we can find a and b such that

$$\begin{gathered} \mathsf{P}\left[\chi^2(T-k) > b\right] = \frac{\alpha}{2}, \\ \mathsf{P}\left[\chi^2(T-k) < a\right] = \frac{\alpha}{2}, \\ \mathsf{P}\left[a \le \chi^2(T-k) \le b\right] = 1-\left(\frac{\alpha}{2}+\frac{\alpha}{2}\right) = 1-\alpha, \end{gathered}$$

which entails that

$$\begin{gathered} \mathsf{P}\left[a \le \frac{(T-k)s^2}{\sigma^2} \le b\right] = 1-\alpha \\ \mathsf{P}\left[\frac{1}{b} \le \frac{\sigma^2}{(T-k)s^2} \le \frac{1}{a}\right] = 1-\alpha \\ \mathsf{P}\left[\frac{(T-k)s^2}{b} \le \sigma^2 \le \frac{(T-k)s^2}{a}\right] = 1-\alpha\ . \end{gathered}$$

It is important to note this is not the smallest confidence interval for σ^2.

6.2. Confidence interval for a linear combination of regression coefficients

Consider now the linear combination $w'\beta$. Then

$$w'\hat{\beta} - w'\beta \sim N\left[0, \sigma^2 w'\left(X'X\right)^{-1}w\right]\ ,$$

hence

$$\frac{w'\hat{\beta} - w'\beta}{\sigma\Delta} \sim N[0,1]$$

where $\Delta = \sqrt{w'(X'X)^{-1}w}$. Since σ is unknown, consider:

$$\begin{aligned}
t &= \frac{w'\hat{\beta} - w'\beta}{s\Delta} \\
&= \frac{w'\hat{\beta} - w'\beta}{\Delta\sigma\sqrt{\frac{s^2}{\sigma^2}}} = \frac{w'\hat{\beta} - w'\beta}{\sigma\Delta} / \sqrt{\frac{(T-k)s^2}{\sigma^2(T-k)}} \\
&= Y / \sqrt{\frac{X}{T-k}}
\end{aligned}$$

where X and Y are independent, $Y \sim N[0,1]$ and $X \sim \chi^2(T-k)$. Thus, t follows a Student t distribution with $T-k$ degrees of freedom:

$$t \sim t(T-k)$$

hence

$$\mathsf{P}\left[-t_{\alpha/2} \leq t(T-k) \leq t_{\alpha/2}\right] = 1-\alpha$$

where $\mathsf{P}\left[t(T-k) > t_{\alpha/2}\right] = \alpha/2$ and

$$\mathsf{P}\left[w'\hat{\beta} - t_{\alpha/2}s\Delta \leq w'\beta \leq w'\hat{\beta} + t_{\alpha/2}s\Delta\right] = 1-\alpha\ .$$

6.3. Confidence region for a regression coefficient vector

We now wish to build a confidence region for a vector $R\beta$ of linear combinations of the elements of β, where $R : q \times k$ and has rank q. Then

$$S(R,\hat{\beta})/\sigma^2 = (R\hat{\beta} - R\beta)'\left(R\left(X'X\right)^{-1}R'\right]^{-1}(R\hat{\beta} - R\beta)/\sigma^2 \sim \chi^2(q)\ .$$

Since σ is unknown, let us consider:

$$F = S(R,\hat{\beta})/qs^2 = \frac{S(R,\hat{\beta})/q\sigma^2}{(T-k)s^2/\sigma^2(T-k)} = \frac{X_1/q}{X_2/(T-k)}$$

where X_1 and X_2 are independent,

$$\begin{aligned}
X_1 = S(R,\hat{\beta})/\sigma^2 \sim \chi^2(q)\ , \\
X_2 = (T-k)s^2/\sigma^2 \sim \chi^2(T-k)\ .
\end{aligned}$$

Thus F follows a Fisher distribution with $(q,\, T-k)$ degrees of freedom:

$$F \sim F(q, T-k)\ .$$

If we define F_α by

$$\mathsf{P}\left[F\left(q, T-k\right) > F_\alpha\right] = \alpha \ ,$$

the set of all vectors $R\beta$ such that $F \leq F_\alpha$:

$$(R\hat{\beta} - R\beta)'\left[R\left(X'X\right)^{-1}R'\right]^{-1}(R\hat{\beta} - R\beta)/qs^2 \leq F_\alpha \ .$$

is a confidence region with level $1-\alpha$ for $R\beta$. This set is a an ellipsoid (*confidence ellipsoid*).

6.4. Prediction intervals

$$y_0 = x_0'\beta + \varepsilon_0$$

where

$$\begin{pmatrix} \varepsilon \\ \varepsilon_0 \end{pmatrix} \sim N\left[0, \sigma^2 I_{T+1}\right] \ .$$

Further

$$\begin{aligned} \hat{y}_0 &= x_0'\hat{\beta} \ , \quad \hat{\beta} = \left(X'X\right)^{-1}X'y, \\ \hat{y}_0 - y_0 &= x_0'(\hat{\beta} - \beta) - \varepsilon_0 \sim N\{0, \sigma^2[1 + x_0'\left(X'X\right)^{-1}x_0]\}. \end{aligned}$$

hence

$$\frac{\hat{y}_0 - y_0}{\sigma\Delta_1} \sim N\left[0, 1\right] \ ,$$

where $\Delta_1 = \left[1 + x_0'\left(X'X\right)^{-1}x_0\right]^{1/2}$, and

$$\frac{\hat{y}_0 - y_0}{s\Delta_1} \sim t\left(T-k\right)$$

where $t_{\alpha/2}$ satisfies

$$\mathsf{P}\left[\hat{y}_0 - t_{\alpha/2}s\Delta_1 \leq y_0 \leq \hat{y}_0 + t_{\alpha/2}s\Delta_1\right] = 1-\alpha \ .$$

6.5. Confidence regions for several predictions

We now consider the problem of predicting a vector of observations y_0 generated according to the same model independently of y :

$$y_0 = X_0\beta + \varepsilon_0 \ ,$$

$$\begin{pmatrix} \varepsilon \\ \varepsilon_0 \end{pmatrix} \sim N\left[0, \sigma^2 I_{T+m}\right] \ ,$$

where X_0 is known but y_0 is not observed. For predicting y_0, let us define:

$$\begin{aligned} \hat{y}_0 &= X_0\hat{\beta}\,, \\ \hat{e}_0 &= y_0 - \hat{y}_0 = \varepsilon_0 - X_0(\hat{\beta} - \beta)\,, \end{aligned}$$

where

$$\begin{aligned} \mathsf{E}(\hat{e}_0) &= 0\,, \\ \mathsf{V}(\hat{e}_0) &= \sigma^2\left[I_m + X_0\left(X'X\right)^{-1}X_0'\right] = \sigma^2 D_0\,, \\ \hat{e}_0 &\sim N\left[0,\, \sigma^2[I_m + X_0\left(X'X\right)^{-1}X_0']\right]\,. \end{aligned}$$

Consequently,

$$\begin{aligned} \hat{e}_0'\mathsf{V}(\hat{e}_0)^{-1}\hat{e}_0 &\sim \chi^2(m)\,, \\ \hat{e}_0'D_0^{-1}\hat{e}_0/\sigma^2 &\sim \chi^2(m)\,. \end{aligned}$$

Since σ^2 is unknown, we replace it by s^2:

$$(T-k)s^2/\sigma^2 \sim \chi^2(T-k)\,.$$

Further, since s^2 is independent of y_0 and $\hat{y}_0 = X\hat{\beta}$, s^2 is independent of $\hat{e}_0$,

$$\begin{aligned} F &= \frac{\hat{e}_0'D_0^{-1}\hat{e}_0}{ms^2} = \frac{\hat{e}_0'D_0^{-1}\hat{e}_0/\sigma^2 m}{(T-k)s^2/\sigma^2(T-k)} \sim F(m,\, T-k)\,, \\ F &= (y_0 - \hat{y}_0)'\left[I_m + X_0\left(X'X\right)^{-1}X_0'\right]^{-1}(y_0 - \hat{y}_0)/ms^2 \sim F(m,\, T-k)\,. \end{aligned}$$

Then the set of vectors y_0 such that

$$F \le F_\alpha(m,\, T-k)$$

is a confidence region for y_0 with level $1-\alpha$.

7. Hypothesis tests

7.0.1 Let us now consider the problem of testing an hypothesis of the form

$$H_0 : w'\beta = w_0 \tag{7.1}$$

where w be a $k \times 1$ vector of constants.To test H_0, it is natural to consider the difference:

$$w'\hat{\beta} - w_0 = w'\left(\hat{\beta} - \beta\right) \sim N\left[0, \sigma^2 w'\left(X'X\right)^{-1}w\right]\,.$$

Under the assumptions of the Gaussian classical linear model, we then have:

$$\begin{aligned}\frac{w'\hat{\beta}-w_0}{\sigma\Delta} &\sim N[0,1] \ ,\Delta=\left[w'\left(X'X\right)^{-1}w\right]^{1/2} \ ,\\ t &= \frac{w'\hat{\beta}-w_0}{s\Delta}\sim t\left(T-k\right) \ .\end{aligned}$$

This suggests the following tests of H_0 :

$$\text{reject } H_0 \text{ at level } \alpha \text{ against } w'\beta-w_0\neq 0 \text{ when } |t|\geq t_{\alpha/2} \quad \text{(two-sided test)} \tag{7.2}$$

$$\text{reject } H_0 \text{ at level } \alpha \text{ against } w'\beta-w_0> 0 \text{ when } t\geq t_{\alpha} \quad \text{(one-sided test)} \tag{7.3}$$

$$\text{reject } H_0 \text{ at level } \alpha \text{ against } w'\beta-w_0< 0 \text{ when } t\leq -t_{\alpha} \quad \text{(one-sided test).} \tag{7.4}$$

An important special case of the above problem consists in testing the value of any given component of β :

$$H_0(\beta_{io}):\beta_i=\beta_{io}$$

where β_i is an element of β.

Let us now consider the more general hypothesis which consists in testing the value of a general vector linear transformation of β :

$$H_0:R\beta=r=\begin{bmatrix} w_1' \\ w_2' \\ \vdots \\ w_q' \end{bmatrix}\beta=\begin{bmatrix} w_1'\beta \\ w_2'\beta \\ \vdots \\ w_q'\beta \end{bmatrix} \tag{7.5}$$

where R is a $q\times k$ fixed matrix with full row rank $[rank(R)=q]$.

7.0.2 Wald-type test. A natural approach then consists in estimating $R\beta$ by $R\hat{\beta}$,and then to examine the difference $R\hat{\beta}-r$. Under H_0,

$$R\hat{\beta}\sim N[r,\Sigma_R] \ , \quad \text{where} \quad \Sigma_R=\sigma^2R\left(X'X\right)^{-1}R' \ .$$

We need a concept of distance between $R\hat{\beta}$ and r. By (5.3),

$$W=(R\hat{\beta}-r)'\Sigma_R^{-1}(R\hat{\beta}-r)\sim\chi^2(q) \quad \text{under } H_0 \ .$$

We tend to reject H_0 when W is too large ($W\geq c$. However, σ^2 and Σ_R are unknown. It is then natural tom replace σ^2 by the estimate s^2, and Σ_R by

$$\hat{\Sigma}_R=s^2R\left(X'X\right)^{-1}R' \ .$$

This yields a Wald-type criterion:

$$
\begin{aligned}
\hat{W} &= (R\hat{\beta}-r)'\hat{\Sigma}_R^{-1}(R\hat{\beta}-r) \\
&= (R\hat{\beta}-r)'\left[s^2 R\left(X'X\right)^{-1}R'\right]^{-1}(R\hat{\beta}-r) \\
&= (R\hat{\beta}-r)'\left[R\left(X'X\right)^{-1}R'\right]^{-1}(R\hat{\beta}-r)/s^2 \\
&= S(R,\hat{\beta})/s^2 \,.
\end{aligned}
$$

Since

$$F = \hat{W}/q = S(R,\hat{\beta})/qs^2 \sim F\left(q, T-k\right)\,,$$

we reject H_0 at level α when

$$F > F_\alpha\left(q, T-k\right)\,. \tag{7.6}$$

7.0.3 Likelihood ratio test. Another approach to testing H_0 consists in looking for a likelihood ratio test. This test is based on focusing on the likelihood function:

$$L\left(y; X\beta, \sigma^2 I_T\right) = \frac{1}{\left(2\pi\sigma^2\right)^{T/2}}\exp\left\{-\frac{1}{2}\frac{(y-X\beta)'(y-X\beta)}{\sigma^2}\right\}\,. \tag{7.7}$$

Let

$$L(\hat{\Omega}) = \max_{\beta,\,\sigma^2} L = \max_{(\beta,\sigma^2)\in\Omega} L \tag{7.8}$$

i.e. we find values of β and σ^2 which maximize "the probability of the observed sample", and

$$L(\hat{\omega}) = \max_{\substack{\beta,\sigma^2 \\ R\beta=r}} L = \max_{(\beta,\sigma^2)\in\omega} L \tag{7.9}$$

i.e. we find values of β and σ^2 which maximize "the probability of the observed sample" and satisfy H_0, where

$$
\begin{gathered}
\Omega = \left\{\left(\beta,\sigma^2\right) : -\infty < \beta_i < +\infty,\ i=1,\ldots,k,\ 0<\sigma^2<+\infty\right\}\,, \\
\omega = \left\{\left(\beta,\sigma^2\right)\in\Omega : R\beta = r\right\}\,.
\end{gathered}
$$

We see easily that

$$0 \le L(\hat{\omega}) \le L(\hat{\Omega})\,,$$

hence

$$
\begin{gathered}
0 \le \frac{L(\hat{\omega})}{L(\hat{\Omega})} \le 1\,, \\
\frac{L(\hat{\Omega})}{L(\hat{\omega})} \ge 1\,.
\end{gathered}
$$

We reject H_0 when

$$LR(y) \equiv \frac{L(\hat{\Omega})}{L(\hat{\omega})} \geq \lambda_\alpha ,$$

where λ_α depends on the level of the test:

$$\mathsf{P}[LR(y) \geq \lambda_\alpha] = \alpha .$$

7.0.4 $L(\hat{\Omega})$ is achieved when $\beta = \hat{\beta}$ and $\sigma^2 = \hat{\sigma}^2$:

$$\begin{aligned} L(\hat{\Omega}) &= \frac{1}{\left(2\pi\hat{\sigma}^2\right)^{T/2}} \exp\left\{-\frac{1}{2}\frac{\left(y - X\hat{\beta}\right)'\left(y - X\hat{\beta}\right)}{\hat{\sigma}^2}\right\} = \frac{1}{\left(2\pi\hat{\sigma}^2\right)^{T/2}} \exp\left\{-\frac{T}{2}\right\} \\ &= \frac{e^{-T/2}}{\left[2\pi\hat{\sigma}^2\right]^{T/2}} = \frac{T^{T/2}e^{-T/2}}{(2\pi)^{T/2}\left[\left(y - X\hat{\beta}\right)'\left(y - X\hat{\beta}\right)\right]^{T/2}} \\ &= \frac{T^{T/2}e^{-T/2}}{(2\pi)^{T/2}S_\Omega^{T/2}} , \end{aligned}$$

where $S_\Omega = \left(y - X\hat{\beta}\right)'\left(y - X\hat{\beta}\right)$.

7.0.5 To find $L(\hat{\omega})$, it is equivalent to maximize

$$\ln(L) = -\frac{T}{2}\ln(2\pi) - \frac{T}{2}\ln\left(\sigma^2\right) - \frac{1}{2\sigma^2}(y - X\beta)'(y - X\beta)$$

under the constraint $R\beta = r$.Consider σ^2 as given. It is then sufficient to solve the problem:

$$\operatorname*{Min}_{\beta}\,(y - X\beta)'(y - X\beta)$$

with restriction $r - R\beta = 0$. Ton do this, we consider the Lagrangian function:

$$\mathscr{L} = (y - X\beta)'(y - X\beta) - \lambda'[r - R\beta] .$$

The optimum $\beta = \tilde{\beta}$ must satisfy the first-order conditions:

$$\frac{\partial \mathscr{L}}{\partial \beta} = -2X'y + 2\left(X'X\right)\tilde{\beta} + R'\lambda = 0 \quad (7.10)$$

$$\frac{\partial \mathscr{L}}{\partial \lambda} = r - R\tilde{\beta} = 0 . \quad (7.11)$$

On multiplying by (7.10) by $R\left(X'X\right)^{-1}$, we get:

$$-2R\left(X'X\right)^{-1}X'y+2R\tilde{\beta}+R\left(X'X\right)^{-1}R'\lambda=0$$
$$R\left(X'X\right)^{-1}R'\lambda=2R\left(X'X\right)^{-1}X'y-2r=2\left[R\hat{\beta}-r\right]$$
$$\lambda=2\left[R\left(X'X\right)^{-1}R'\right]^{-1}\left[R\hat{\beta}-r\right]\ .$$

By (7.10),

$$2\left(X'X\right)\tilde{\beta} = 2X'y-R'\lambda \tag{7.12}$$
$$= 2X'y-2R'\left[R\left(X'X\right)^{-1}R'\right]^{-1}\left[R\hat{\beta}-r\right] \tag{7.13}$$

hence

$$\begin{aligned}\tilde{\beta} &= \left(X'X\right)^{-1}X'y-\left(X'X\right)^{-1}R'\left[R\left(X'X\right)^{-1}R'\right]^{-1}\left[R\hat{\beta}-r\right]\\ &= \hat{\beta}+\left(X'X\right)^{-1}R'\left[R\left(X'X\right)^{-1}R'\right]^{-1}\left[r-R\hat{\beta}\right]\ .\end{aligned}$$

We see that $\tilde{\beta}$ does not depend on σ^2. Substituting $\tilde{\beta}$ in $\ln(L)$, we see that

$$\ln(L)=-\frac{T}{2}\ln(2\pi)-\frac{T}{2}\ln\sigma^2-\frac{1}{2\sigma^2}S_\omega$$

where $S_\omega=\left(y-X\tilde{\beta}\right)'\left(y-X\tilde{\beta}\right)$, from which we get

$$\frac{\partial\ln(L)}{\partial\sigma^2}=-\frac{T}{2\sigma^2}+\frac{S_\omega}{2\sigma^4}=0$$

at the optimum, hence

$$\tilde{\sigma}^2=S_\omega/T=\left(y-X\tilde{\beta}\right)'\left(y-X\tilde{\beta}\right)/T\ ,$$
$$L(\hat{\omega})=\frac{T^{T/2}e^{-T/2}}{(2\pi)^{T/2}S_\omega^{T/2}}\ ,$$

The likelihood ratio test is given by the critical region:

$$\frac{L(\hat{\Omega})}{L(\hat{\omega})}=\left(\frac{S_\omega}{S_\Omega}\right)^{T/2}\geq\lambda_\alpha$$

or, equivalently,

$$\frac{S_\omega}{S_\Omega}\geq\lambda_\alpha^{2/T}\ . \tag{7.14}$$

Since

$$\begin{aligned} S_\omega &= (y - X\tilde{\beta})'(y - X\tilde{\beta}) \\ &= (y - X\hat{\beta})'(y - X\hat{\beta}) + (\hat{\beta} - \tilde{\beta})'\left(X'X\right)(\hat{\beta} - \tilde{\beta}) \\ &= S_\Omega + (\hat{\beta} - \tilde{\beta})'\left(X'X\right)(\hat{\beta} - \tilde{\beta})\ , \end{aligned}$$

we also see that

$$\begin{aligned} S_\omega - S_\Omega &= \left(r - R\hat{\beta}\right)'\left[R\left(X'X\right)^{-1}R'\right]^{-1}R\left(X'X\right)^{-1}\left(X'X\right)\left(X'X\right)^{-1} \\ &\quad R'\left[R\left(X'X\right)^{-1}R'\right]^{-1}\left[r - R\hat{\beta}\right] \\ &= \left(r - R\hat{\beta}\right)'\left[R\left(X'X\right)^{-1}R'\right]^{-1}\left[r - R\hat{\beta}\right] \\ &= (R\hat{\beta} - r)'\left[R\left(X'X\right)^{-1}R'\right]^{-1}(R\hat{\beta} - r) = S(R, \hat{\beta}) \\ &= \left(qs^2\right)F\ , \end{aligned}$$

hence

$$F = \frac{S_\omega - S_\Omega}{qs^2} = \frac{(S_\omega - S_\Omega)/q}{S_\Omega/(T-k)}$$

and

$$\begin{gathered} \frac{S_\omega}{S_\Omega} = \frac{S_\Omega + \left(qs^2\right)F}{S_\Omega} = 1 + \frac{\left(qs^2\right)F}{(T-k)s^2} = 1 + \frac{q}{T-k}F \geq \lambda_\alpha^{2/T} \\ \Longleftrightarrow F \geq \frac{T-k}{q}\left(\lambda_\alpha^{2/T} - 1\right) = F_\alpha\ . \end{gathered}$$

The likelihood ratio test of $H_0 : R\beta = r$ has the critical region

$$F \equiv \frac{(S_\omega - S_\Omega)/q}{S_\Omega/(T-k)} \geq F_\alpha(q, T-k)$$

where

$$F \sim F(q, T-k)\ .$$

This is an easy method for testing $H_0 : R\beta = r$. Note also that:

$$\begin{aligned} LR &= \left(\frac{S_\omega}{S_\Omega}\right)^{T/2} = \left(1 + \frac{q}{T-k}F\right)^{T/2}\ , \\ F &= \frac{T-k}{q}\left(LR^{2/T} - 1\right)\ . \end{aligned}$$

8. Estimator optimal properties with Gaussian errors

When errors are Gaussian, the OLS estimators $\hat{\beta}_i, i = 1, \ldots, k$ and $s^2 = \left(y - X\hat{\beta}\right)' \left(y - X\hat{\beta}\right) / (T - k)$ have minimum variance in the class of all unbiased estimators of β_i, $i = 1, \ldots, k$, and σ^2 respectively [see Rao (1973, section 5a)].

References

Maddala, G. S. (1977), *Econometrics*, McGraw-Hill, New York.

Rao, C. R. (1973), *Linear Statistical Inference and its Applications*, second edn, John Wiley & Sons, New York.

Schmidt, P. (1976), *Econometrics*, Marcel Dekker, New York.

Theil, H. (1971), *Principles of Econometrics*, John Wiley & Sons, New York.