Chapter 2

Inferences in Regression Analysis

We will focus on the normal error model:

$$Y_i = \beta_0 + \beta_1 X_i + \epsilon_i,$$

where

1. β_0 and β_1 are parameters to be estimated,
2. X_i's are known constants,
3. ϵ_i's are *iid* (independent and identically distributed) $N(0, \sigma^2)$.

For convenience, we often use b_0 and b_1 as the estimators of β_0 and β_1 respectively, *i.e.*, $b_0 = \hat{\beta}_0$ and $b_1 = \hat{\beta}_1$.

2.1 Inferences concerning $\hat{\beta}_1$

2.1.1 Distribution of $\hat{\beta}_1$

We want to test

$$\boxed{H_0 : \beta_1 = 0 \quad \text{and} \quad H_1 : \beta_1 \neq 0.}$$

For the hypothesis test above, we have to know the distribution of $\hat{\beta}_1$ under H_0.

Fact. *If the independent random variable Y_i has a normal distribution with $N(\mu_i, \sigma_i^2)$, then the linear combination $\sum_{i=1}^n (c_i Y_i + d_i)$ also has a normal with $N(\sum_{i=1}^n (c_i \mu_i + d_i), \sum_{i=1}^n c_i^2 \sigma_i^2)$.*

We derived the regression estimator

$$\hat{\beta}_1 = \frac{S_{xy}}{S_{xx}} = \frac{\sum_{i=1}^n (X_i - \bar{X})(Y_i - \bar{Y})}{\sum_{i=1}^n (X_i - \bar{X})^2} = \sum_{i=1}^n k_i (Y_i - \bar{Y}) = \sum_{i=1}^n k_i Y_i, \tag{2.1}$$

where $k_i = (X_i - \bar{X})/S_{xx}$ and $S_{xx} = \sum_{i=1}^n (X_i - \bar{X})^2$. Then we have

$$\hat{\beta}_1 \sim N\Big(\sum_{i=1}^n k_i(\beta_0 + \beta_1 X_i), \sum_{i=1}^n k_i^2 \sigma^2\Big).$$

Using $\sum_{i=1}^n k_i = 0$, $\sum_{i=1}^n k_i X_i = 1$, $\sum_{i=1}^n k_i^2 = 1/S_{xx}$, we have

$$\boxed{\hat{\beta}_1 \sim N\Big(\beta_1, \frac{\sigma^2}{S_{xx}}\Big).}$$

Theorem 2.1. *The regression estimator $\hat{\beta}_1$ is a BLUE (best linear unbiased estimator) of β_1. The term "best" here is used in the sense of minimum variance.*

Proof. Let $\hat{\beta}_1^*$ be another unbiased linear estimator of the form:

$$\hat{\beta}_1^* = \sum_{i=1}^n c_i Y_i,$$

where the c_i's are constants. Since $\hat{\beta}_1^*$ is unbiased, it should satisfy

$$E(\hat{\beta}_1^*) = \sum_{i=1}^n c_i E(Y_i) = \sum_{i=1}^n c_i(\beta_0 + \beta_1 X_i) = \beta_0 \sum_{i=1}^n c_i + \beta_1 \sum_{i=1}^n c_i X_i = \beta_1.$$

Hence, we have the following two conditions for the BLUE of β_1:

$$\boxed{\sum_{i=1}^n c_i = 0 \quad \text{and} \quad \sum_{i=1}^n c_i X_i = 1.} \tag{2.2}$$

The variance of $\hat{\beta}_1^*$ is

$$\text{Var}(\hat{\beta}_1^*) = \sum_{i=1}^n c_i^2 \,\text{Var}(Y_i) = \sigma^2 \sum_{i=1}^n c_i^2. \tag{2.3}$$

We know $\text{Var}(\hat{\beta}_1^* - \hat{\beta}_1) \geq 0$. Let's try to find $\text{Var}(\hat{\beta}_1^* - \hat{\beta}_1)$. Then we have

$$\text{Var}(\hat{\beta}_1^* - \hat{\beta}_1) = \text{Var}\left(\sum_{i=1}^n (c_i - k_i) Y_i\right) = \sigma^2 \sum_{i=1}^n (c_i - k_i)^2 = \sigma^2 \sum_{i=1}^n (c_i^2 + k_i^2 - 2c_i k_i).$$

Note that $\sum_{i=1}^n c_i k_i = \sum_{i=1}^n c_i(X_i - \bar{X})/S_{xx} = (\sum c_i X_i - \sum c_i \bar{X})/S_{xx} = 1/S_{xx}$ and $\sum_{i=1}^n k_i^2 = \sum_{i=1}^n [(X_i - \bar{X})/S_{xx}]^2 = 1/S_{xx}$. Hence, we have $\sum_{i=1}^n c_i k_i = \sum_{i=1}^n k_i^2$. It is immediate from $\text{Var}(\hat{\beta}_1^* - \hat{\beta}_1) \geq 0$ that

$$\text{Var}(\hat{\beta}_1^* - \hat{\beta}_1) = \sigma^2 \sum_{i=1}^n (c_i^2 - k_i^2) = \text{Var}(\hat{\beta}_1^*) - \text{Var}(\hat{\beta}_1) \geq 0.$$

Hence, $\text{Var}(\hat{\beta}_1^*) \geq \text{Var}(\hat{\beta}_1)$. So $\hat{\beta}_1$ has minimum variance among all unbiased linear estimators. $\square$

2.1.2 Sampling distribution of $(\hat{\beta}_1 - \beta_1)/\text{SE}(\hat{\beta}_1)$

Since $\hat{\beta}_1 \sim N(\beta_1, \sigma^2/S_{xx})$, we have

$$\frac{\hat{\beta}_1 - \beta_1}{\sqrt{\sigma^2/S_{xx}}} \sim N(0, 1).$$

We need to estimate σ^2, we will use MSE as an estimator of σ^2:

$$\text{MSE} = \frac{1}{n-2}\sum_{i=1}^{n}(Y_i - \hat{Y}_i)^2.$$

Note that

$$\frac{(n-2)\text{MSE}}{\sigma^2} \sim \chi^2_{n-2} \quad \text{and} \quad \frac{N(0,1)}{\sqrt{\chi^2_d/d}} \sim t(d).$$

Hence, we have

$$\frac{\hat{\beta}_1 - \beta_1}{\sqrt{\sigma^2/S_{xx}}} \Big/ \sqrt{\frac{(n-2)\text{MSE}}{\sigma^2}\Big/(n-2)} \sim t(n-2),$$

and it follows that

$$\boxed{\frac{\hat{\beta}_1 - \beta_1}{\sqrt{\text{MSE}/S_{xx}}} = \frac{\hat{\beta}_1 - \beta_1}{\text{SE}(\hat{\beta}_1)} \sim t(n-2)}$$

where $\text{SE}(\hat{\beta}_1) = \sqrt{\text{MSE}/S_{xx}}$.

2.1.3 Confidence interval for β_1

Since $(\hat{\beta}_1 - \beta_1)/\text{SE}(\hat{\beta}_1) \sim t(n-2)$, we have

$$P\Big[t(\tfrac{\alpha}{2}; n-2) \le \tfrac{\hat{\beta}_1 - \beta_1}{\text{SE}(\hat{\beta}_1)} \le t(1-\tfrac{\alpha}{2}; n-2)\Big] = 1-\alpha.$$

Using the symmetry of t distribution about 0, we have $t(\frac{\alpha}{2}; n-2) = -t(1-\frac{\alpha}{2}; n-2)$.

Hence, the $1-\alpha$ confidence limits for β_1 are

$$\boxed{\hat{\beta}_1 \pm t(1-\tfrac{\alpha}{2}; n-2)\cdot \text{SE}(\hat{\beta}_1).}$$

2.1.4 Tests for β_1

- **Two-sided test**: $H_0 : \beta_1 = 0$ vs. $H_1 : \beta_1 \neq 0$.

  Test statistic:

$$T = \frac{\hat{\beta}_1}{\text{SE}(\hat{\beta}_1)}$$

Decision rule (at the significance level α):

If $|T| \leq t(1-\frac{\alpha}{2}; n-2)$, accept H_0.

If $|T| > t(1-\frac{\alpha}{2}; n-2)$, reject H_0.

- **One-sided test**: $H_0 : \beta_1 \leq 0$ vs. $H_1 : \beta_1 > 0$.

  Test statistic:

$$T = \frac{\hat{\beta}_1}{\mathrm{SE}(\hat{\beta}_1)}$$

  Decision rule (at the significance level α):

  If $T \leq t(1-\alpha; n-2)$, accept H_0.

  If $T > t(1-\alpha; n-2)$, reject H_0.

- **Test when** $H_0 : \beta_1 = \beta_{10}$ **or** $H_0 : \beta_1 < \beta_{10}$

  Test statistic:

$$T = \frac{\hat{\beta}_1 - \beta_{10}}{\mathrm{SE}(\hat{\beta}_1)}$$

  Decision rule (at the significance level α): the same as the above.

2.1.5 A geometry of the BLUE of β_1

The construction of BLUE of β_1 can be explained by using the geometric description as in the figure. The goal is to minimize the variance, $\sigma^2 \sum_{i=1}^n c_i^2$, in (2.3) together with the conditions, $\sum_{i=1}^n c_i = 0$ and $\sum_{i=1}^n c_i X_i = 1$, in (2.2). This is equivalent to finding the minimization of $\sum_{i=1}^n c_i^2$ (hyper-sphere in $\mathbb{R}^n$) with the two constraints (two hyper-planes in $\mathbb{R}^n$).

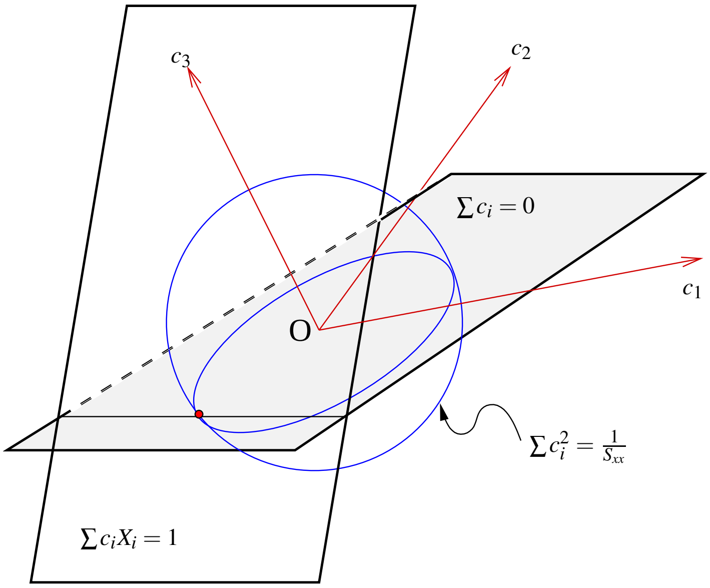

Figure 2.1: Geometry of the BLUE of β_1 with $n = 3$.

2.1.6 BLUE: Optimization with the constraints

We are to minimize $\sum_{i=1}^n c_i^2$ subject to $\sum_{i=1}^n c_i = 0$ and $\sum_{i=1}^n c_i X_i = 1$. Thus, this minimization can be obtained by using Lagrange multipliers. The auxiliary function with Lagrange multipliers (λ_1 and λ_2) is

$$\Psi = \sum_{i=1}^n c_i^2 + \lambda_1 \sum_{i=1}^n c_i + \lambda_2 \Big(\sum_{i=1}^n c_i X_i - 1 \Big).$$

It follows from $\partial\Psi/\partial c_i = 0$ that

$$2c_i + \lambda_1 X_i + \lambda_2 = 0. \tag{2.4}$$

Taking the sum of the above, we have

$$2\sum_{i=1}^n c_i + \lambda_1 n\bar{X} + n\lambda_2 = 0.$$

It is immediate from the condition $\sum_{i=1}^n c_i = 0$ that

$$\lambda_1 n\bar{X} + n\lambda_2 = 0. \tag{2.5}$$

Next, multiplying c_i to (2.4) gives

$$2c_i^2 + \lambda_1 c_i X_i + \lambda_2 c_i = 0.$$

Taking the sum of the above, we have

$$2\sum_{i=1}^n c_i^2 + \lambda_1 \sum_{i=1}^n c_i X_i + \lambda_2 \sum_{i=1}^n c_i = 0.$$

It is immediate from the two conditions in (2.2) that

$$2\sum_{i=1}^n c_i^2 + \lambda_1 = 0. \tag{2.6}$$

Solving (2.5) and (2.6) for λ_1 and λ_2, we have

$$\lambda_1 = -2\sum_{i=1}^n c_i^2 \quad \text{and} \quad \lambda_2 = 2\sum_{i=1}^n c_i^2 \bar{X}.$$

Substituting the above results into (2.4) gives

$$c_i = (X_i - \bar{X}) \sum_{i=1}^n c_i^2. \tag{2.7}$$

Taking the square of (2.7) gives

$$c_i^2 = (X_i - \bar{X})^2 \Big(\sum_{i=1}^n c_i^2 \Big)^2.$$

After taking the sum of the above, we can solve for $\sum_{i=1}^n c_i^2$ which results in

$$\sum_{i=1}^n c_i^2 = \frac{1}{\sum_{i=1}^n (X_i - \bar{X})^2} = \frac{1}{S_{xx}}. \tag{2.8}$$

It is immediate upon substituting (2.8) into (2.7) that

$$c_i = \frac{X_i - \bar{X}}{S_{xx}}.$$

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2.2 Inferences concerning β_0

The estimator for β_0 is

$$\hat{\beta}_0 = \bar{Y} - \hat{\beta}_1 \bar{X}$$

The expectation and variance of $\hat{\beta}_0$ can be shown as follows:

$$E(\hat{\beta}_0) = \beta_0$$
$$\mathrm{Var}(\hat{\beta}_0) = \sigma^2 \Big[\frac{1}{n} + \frac{\bar{X}^2}{S_{xx}}\Big].$$

Let $k_i = (X_i - \bar{X})/S_{xx}$. We have studied $\hat{\beta}_1 = \sum k_i Y_i$ from (2.1). Thus, we have

$$\hat{\beta}_0 = \bar{Y} - \hat{\beta}_1 \bar{X} = \sum_{i=1}^{n} \frac{1}{n} Y_i - \sum_{i=1}^{n} k_i Y_i \bar{X} = \sum_{i=1}^{n} \Big(\frac{1}{n} - k_i \bar{X}\Big) Y_i. \tag{2.9}$$

Note that ϵ_i are independent, so is Y_i. Hence, we have

$$\begin{aligned}
\mathrm{Var}(\hat{\beta}_0) &= \sum_{i=1}^{n} \Big(\frac{1}{n} - k_i \bar{X}\Big)^2 \mathrm{Var}(Y_i) \\
&= \sum_{i=1}^{n} \Big(\frac{1}{n^2} - \frac{2}{n} k_i \bar{X} + k_i^2 \bar{X}^2\Big) \sigma^2 \\
&= \Big(\frac{1}{n} + \frac{\bar{X}^2}{S_{xx}}\Big) \sigma^2.
\end{aligned}$$

Hence, an estimator of $\mathrm{Var}(\hat{\beta}_0)$, also denoted by $[\mathrm{SE}(\hat{\beta}_0)]^2$, is obtained by replacing σ^2 by MSE:

$$\widehat{\mathrm{Var}}(\hat{\beta}_0) = [\mathrm{SE}(\hat{\beta}_0)]^2 = \mathrm{MSE}\Big[\frac{1}{n} + \frac{\bar{X}^2}{S_{xx}}\Big].$$

Similar to the case of $\hat{\beta}_1$, we have the following test statistic:

$$\frac{\hat{\beta}_0 - \beta_0}{\mathrm{SE}(\hat{\beta}_0)} \sim t(n-2).$$

The confidence limits: $\hat{\beta}_0 \pm t(1 - \frac{\alpha}{2}; n-2) \cdot \mathrm{SE}(\hat{\beta}_0)$.

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2.3 Interval estimation of $\mu_{Y_h} = E(Y_h) = \beta_0 + \beta_1 X_h$

Let X_h denote the level of X for which we wish to estimate the mean response, $E(Y_h) = \beta_0 + \beta_1 X_h$. We want to draw inference about $E(Y_h)$

Given the estimators $\hat{\beta}_0$ and $\hat{\beta}_1$ of the regression parameters in the regression function $\mu_Y = E(Y) = \beta_0 + \beta_1 X$, we estimate the regression function as

$$\hat{\mu}_Y = \hat{Y} = \hat{\beta}_0 + \hat{\beta}_1 X.$$

Thus, when $X = X_h$, the estimator of $\mu_{Y_h} = E(Y_h)$ denoted by $\hat{Y}_h$ is given by

$$\hat{Y}_h = \hat{\beta}_0 + \hat{\beta}_1 X_h.$$

The estimators $\hat{\beta}_0$ and $\hat{\beta}_1$ all consist of linear combination of Y_i's. Hence, $\hat{Y}_h = \hat{\beta}_0 + \hat{\beta}_1 X_h$ is also a linear combination of Y_i and so $\hat{\beta}_0 + \hat{\beta}_1 X_h$ has a normal distribution. To find the parameters for this normal distribution, we need to find $E(\hat{Y}_h)$ and $\mathrm{Var}(\hat{Y}_h)$. First, $E(\hat{Y}_h)$ is obtained by

$$E(\hat{Y}_h) = E(\hat{\beta}_0 + \hat{\beta}_1 X_h) = \beta_0 + \beta_1 X_h = \mu_{Y_h}.$$

Next, using $\hat{\beta}_1 = \sum_{i=1}^n k_i Y_i$ from (2.1) and $\hat{\beta}_0 = \sum_{i=1}^n \left(\frac{1}{n} - k_i \bar{X}\right) Y_i$ from (2.9), we have

$$\hat{Y}_h = \hat{\beta}_0 + \hat{\beta}_1 X_h = \sum_{i=1}^{n} \left\{ \frac{1}{n} + (X_h - \bar{X}) k_i \right\} Y_i.$$

It is immediate that

$$\begin{aligned}
\mathrm{Var}(\hat{Y}_h) &= \mathrm{Var}\Big[\sum_{i=1}^{n}\Big\{\frac{1}{n} + (X_h - \bar{X})k_i\Big\}Y_i\Big] \\
&= \sum_{i=1}^{n}\Big\{\frac{1}{n} + (X_h - \bar{X})k_i\Big\}^2 \mathrm{Var}(Y_i) \\
&= \sum_{i=1}^{n}\Big\{\frac{1}{n^2} + \frac{2}{n}(X_h - \bar{X})k_i + (X_h - \bar{X})^2 k_i^2\Big\}\sigma^2 \\
&= \Big\{\frac{1}{n} + \frac{(X_h - \bar{X})^2}{S_{xx}}\Big\}\sigma^2,
\end{aligned}$$

and

$$\widehat{\mathrm{Var}}(\hat{Y}_h) = [\mathrm{SE}(\hat{Y}_h)]^2 = \mathrm{MSE}\Big[\frac{1}{n} + \frac{(X_h - \bar{X})^2}{S_{xx}}\Big].$$

Notice that $\mathrm{SE}(\hat{Y}_h)$ is minimized when $X_h = \bar{X}$.

Thus, we have the distribution

$$Z = \frac{\hat{Y}_h - E(\hat{Y}_h)}{\sqrt{\mathrm{Var}(\hat{Y}_h)}} = \frac{\hat{Y}_h - \mu_{Y_h}}{\sqrt{\mathrm{Var}(\hat{Y}_h)}} \;\sim\; N(0,1).$$

Replacing $\sqrt{\mathrm{Var}(\hat{Y}_h)}$ with $\sqrt{\widehat{\mathrm{Var}}(\hat{Y}_h)} = \mathrm{SE}(\hat{Y}_h)$, we have the following test statistic

$$T = \frac{\hat{Y}_h - E(\hat{Y}_h)}{\sqrt{\widehat{\mathrm{Var}}(\hat{Y}_h)}} = \frac{\hat{Y}_h - \mu_{Y_h}}{\mathrm{SE}(\hat{Y}_h)} \;\sim\; t(n-2).$$

Converting the above for μ_{Y_h}, we have the confidence limits for μ_{Y_h}:

$$\hat{Y}_h \pm t(1 - \frac{\alpha}{2}; n-2) \cdot \mathrm{SE}(\hat{Y}_h).$$

That is,

$$\hat{Y}_h \pm t(1 - \frac{\alpha}{2}; n-2) \cdot \sqrt{\mathrm{MSE} \cdot \Big\{\frac{1}{n} + \frac{(X_h - \bar{X})^2}{S_{xx}}\Big\}}.$$

2.4 Prediction Interval of *New* observation, $Y_{h(\text{new})}$

There is often confusion regarding the implication of the word *prediction*. Obviously the statistic $\hat{Y}_h = \hat{\beta}_0 + \hat{\beta}_1 X_h$, the point on the regression line at $X = X_h$, serves the dual purpose as the estimate of mean response ($E(Y_h)$ at $X = X_h$) and the predicted value. However, it is not appropriate for establishing any form of inference on a *future* single observation. Suppose the *mean response* at a fixed $X = X_h$ is not of interest. Rather, one is interested in some type of bound on a *future* single response observation at $X = X_h$.

Consider a *future* single observation at $X = X_h$, denoted symbolically by $Y_{h(\text{new})}$, independent of $\hat{Y}_h = \hat{\beta}_0 + \hat{\beta}_1 X_h$. We can standardize by considering

$$\boxed{\text{Var}(Y_{h(\text{new})} - \hat{Y}_h) = \text{Var}(Y_{h(\text{new})}) + \text{Var}(\hat{Y}_h) = \sigma^2\Big\{1 + \frac{1}{n} + \frac{(X_h - \bar{X})^2}{S_{xx}}\Big\}.}$$

It immediate from $E(Y_{h(\text{new})} - \hat{Y}_h) = 0$ that

$$Z = \frac{Y_{h(\text{new})} - \hat{Y}_h - 0}{\sigma\sqrt{1 + \frac{1}{n} + \frac{(X_h - \bar{X})^2}{S_{xx}}}} \ \sim \ N(0, 1).$$

Replacing σ^2 with MSE, we have

$$T = \frac{Y_{h(\text{new})} - \hat{Y}_h}{\sqrt{\text{MSE}} \cdot \sqrt{1 + \frac{1}{n} + \frac{(X_h - \bar{X})^2}{S_{xx}}}} \ \sim \ t(n-2).$$

Converting the above for $Y_{h(\text{new})}$, we have the prediction limits for μ_{Y_h}:

$$\hat{Y}_h \pm t(1 - \frac{\alpha}{2}; n-2) \cdot \sqrt{\text{MSE} \cdot \Big\{1 + \frac{1}{n} + \frac{(X_h - \bar{X})^2}{S_{xx}}\Big\}}.$$

These limits are the same as

$$\hat{Y}_h \pm t(1 - \frac{\alpha}{2}; n-2) \cdot \sqrt{\text{MSE} + \text{MSE} \cdot \Big\{\frac{1}{n} + \frac{(X_h - \bar{X})^2}{S_{xx}}\Big\}}$$

$$\hat{Y}_h \pm t(1 - \frac{\alpha}{2}; n-2) \cdot \sqrt{\text{MSE} + \widehat{\text{Var}}(\hat{Y}_h)}.$$

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2.5 ANOVA approach to regression analysis

2.5.1 Partition of total sum of squares

In any regression, the analyst will observe variation in the response. We want to look at the variation of Y_i and $\hat{Y}_i$. Note that $\sum Y_i = \sum \hat{Y}_i$, *i.e.*, $\bar{Y} = \bar{\hat{Y}}$.

$$\underbrace{Y_i - \bar{Y}}_{\substack{\text{deviation}\\ \text{of } Y_i}} = \underbrace{\hat{Y}_i - \bar{Y}}_{\substack{\text{deviation}\\ \text{of } \hat{Y}_i}} + \underbrace{Y_i - \hat{Y}_i}_{\hat{\epsilon}_i}$$

Then the sum of squares become

$$\begin{aligned}\sum_{i=1}^{n}(Y_i - \bar{Y})^2 &= \sum_{i=1}^{n}(\hat{Y}_i - \bar{Y} + Y_i - \hat{Y}_i)^2 \\ &= \sum(\hat{Y}_i - \bar{Y})^2 + \sum(Y_i - \hat{Y}_i)^2 + 2\sum(\hat{Y}_i - \bar{Y})(Y_i - \hat{Y}_i)\end{aligned}$$

Considering the last term,

$$\begin{aligned}2\sum(\hat{Y}_i - \bar{Y})(Y_i - \hat{Y}_i) &= 2\sum \hat{Y}_i(Y_i - \hat{Y}_i) - 2\sum \bar{Y}(Y_i - \hat{Y}_i) \\ &= 2\sum \hat{Y}_i\hat{\epsilon}_i - 2\sum \bar{Y}\hat{\epsilon}_i = 0,\end{aligned}$$

We have

$$\sum_{i=1}^{n}(Y_i - \bar{Y})^2 = \sum(\hat{Y}_i - \bar{Y})^2 + \sum(Y_i - \hat{Y}_i)^2.$$

We denote this partition as follows:

$$\text{SSTO} = \text{SSR} + \text{SSE}.$$

The coefficient of determination (R^2) is defined as

$$\boxed{R^2 = \frac{\text{SSR}}{\text{SSTO}} = 1 - \frac{\text{SSE}}{\text{SSTO}}.}$$

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Table 2.1: ANOVA Table:

Source	df	SS	MS	F
Regression	1	SSR	MSR=SSR	$F = \frac{\text{MSR}}{\text{MSE}}$
Error	$n-2$	SSE	$\text{MSE} = \text{SSE}/(n-2)$	
Total	$n-1$	SSTO		

This R^2 tells us what proportion of your variation in Y is explained by the regression on X. Clearly $0 \le R^2 \le 1$.

Notice that the df decomposition: $n-1 \;=\; 1 \;+\; (n-2)$.

Why SSR has only df=1? Plugging $\hat{Y}_i = \hat{\beta}_0 + \hat{\beta}_1 X_i$ and $\bar{Y} = \hat{\beta}_0 + \hat{\beta}_1 \bar{X}$ into $\text{SSR} = \sum(\hat{Y}_i - \bar{Y})^2$, we have

$$\text{SSR} = \text{MSR} = \hat{\beta}_1^2 \cdot \sum(X_i - \bar{X})^2 = \hat{\beta}_1^2 \cdot S_{xx}.$$

Here SSR has only one squared quantity ($\hat{\beta}_1^2$).

2.5.2 F test

Fact. *If the errors ϵ_i have a* iid $N(0, \sigma^2)$, *then SSE and SSR are independent and have scaled χ^2 distribution.*

Thus, we have

1. $\text{SSE}/\sigma^2 \;\sim\; \chi^2_{n-2}$.

2. $\text{SSR}/\sigma^2 \;\sim\; \chi^2_1$ *under* $H_0 : \beta_1 = 0$.

If $\beta_1 \neq 0$, then SSR/σ^2 *follows a non-central χ^2_1 and is usually larger than a χ^2_1.*

Fact. *If $U \sim \chi^2_a$ and $V \sim \chi^2_b$, and U, V are independent. then*

$$\frac{U/a}{V/b} \sim F(a, b).$$

Hence, if the errors are $N(0, \sigma^2)$, then under $H_0 : \beta_1 = 0$

$$F^* = \frac{\text{SSR}/\sigma^2}{\text{SSE}/[(n-2)\sigma^2]} = \frac{\text{MSR}}{\text{MSE}} \sim F(1, n-2)$$

Decision rule:

If $F^* \leq F(1-\alpha; 1, n-2)$, accept $H_0 : \beta_1 = 0$.

If $F^* > F(1-\alpha; 1, n-2)$, reject $H_0 : \beta_1 = 0$.

2.6 Case when X is random

In all of the preceding development, we assume throughout that X_i is a known constant. All the theories developed are based on this assumption. But two other situations occur frequently in practice, however:

1. The variables X and Y are both random and are observations from a joint density function.

   $\Rightarrow$ In this case, generally less interest is associated with prediction than in the case when X is non-random. Correlation analysis is more suitable

2. The variable X is a measure with non-ignorable error

   $\Rightarrow$ Use EIV (Errors-In-Variables) model.

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