



Statistics: Regression

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All models are wrong, but some are useful.

Contents

Chapter 1 Review of statistics	1
1.1 Random Vectors	1
1.2 Affine Transformation	1
Chapter 2 Regression Analysis (SLR)	2
2.1 Simple Linear Regression	2
2.1.1 Assumptions	2
2.1.2 Least Squares Estimations	3
2.1.3 Fitted Values & Residuals	4
2.1.4 Degree of freedom	4
2.1.5 (Sample) Error variance	4
2.2 Goodness of Fit: R -square	5
2.2.1 TSS, RSS, FSS	5
2.2.2 Cofficient of Determination(R^2)	5
2.3 Affine Transformations	5
2.3.1 $\tilde{y}_i = ay_i + b$	6
2.3.2 $\tilde{x}_i = ax_i + b$	6
2.3.3 Regress x on y instead	6
2.4 Regression Through the Origin	6
2.5 LS Estimators Properties	7
2.5.1 Unbiasedness of LS Estimators $E(\hat{\beta}_1) = \beta_1, E(\hat{\beta}_0) = \beta_0$	7
2.5.2 Mean squared error(MSE) of LS Estimators = $Var(\hat{\beta}_1) = \sigma^2 \frac{1}{S_{xx}}, Var(\hat{\beta}_0) = \sigma^2 \left(\frac{1}{n} + \frac{\bar{x}^2}{S_{xx}} \right)$	7
2.6 Normality Assumption	8
2.7 Distribution of LS Estimators	8
2.8 Hypothesis Testing (T-test)	8
2.8.1 Testing for the Slope	8
2.8.2 Testing for the Intercept	9
2.9 ANOVA Table & F-Test	9

2.9.1	Degrees of Freedom	9
2.9.2	ANOVA Table	9
2.9.3	F-Test (equivalent to t-test)	10
2.10	Estimation and Prediction	10
2.10.1	Estimation (always reported for a parameter: $\beta_0 + \beta_1 x^* = \mathbb{E}(Y x^*)$)	10
2.10.2	Prediction (is reported for the value of a random variable Y^*)	11
2.10.3	Simultaneous Confidence	11
2.10.4	Confidence Band (Larger than CI)	12
2.11	Maximum likelihood estimators with normal error terms	13
Chapter 3	Multiple Linear Regression (MLR): Basic	14
3.1	Basic	14
3.1.1	Assumptions of errors	14
3.1.2	Matrix Representation	14
3.2	Parameter Estimation $\hat{\beta}$	15
3.2.1	Parameter Estimation $\hat{\beta} = (\mathbf{X}^T \mathbf{X})^{-1} \mathbf{X}^T y$	15
3.2.2	Hat Matrix $\mathbf{H}_{n \times n} = \mathbf{X} (\mathbf{X}^T \mathbf{X})^{-1} \mathbf{X}^T$	15
3.3	Residuals	16
3.3.1	Residuals: $\mathbf{r} = (\mathbf{I} - \mathbf{H})y$	16
3.3.2	Error Variance: $\hat{\sigma}^2 = \frac{\mathbf{r}^T \mathbf{r}}{n-p}$	16
3.4	Properties of Residuals	16
3.4.1	Orthogonal to column of data $\mathbf{X}^T \mathbf{e}_j = 0$	16
3.4.2	Orthogonal to fitted value $\hat{y}^T \mathbf{e}_j = 0$	16
3.5	Properties of \mathbf{H}	17
3.5.1	$\mathbf{H}\mathbf{X} = \mathbf{X}$	17
3.5.2	Symmetric: $\mathbf{H}^T = \mathbf{H}$	17
3.5.3	Idempotent: $\mathbf{H}\mathbf{H} = \mathbf{H}\mathbf{H}^T = \mathbf{H}^T\mathbf{H} = \mathbf{H}$	17
3.5.4	$\mathbf{H}(\mathbf{I} - \mathbf{H}) = \mathbf{0}$	17
3.5.5	$(\mathbf{I} - \mathbf{H})$ is also symmetric and idempotent	17
3.5.6	$\text{trace}(\mathbf{H}) = p$	17
3.6	Geometric Representation of LS	17
3.6.1	Estimation Space	17
3.6.2	Error Space	18

3.7	Coefficient of determination, R-Square	18
3.8	Properties of LS Estimators	19
3.8.1	Unbiased: $\mathbb{E}(\hat{\beta}) = \beta$	19
3.8.2	Variance-Covariance Matrix of $\hat{\beta}$: $\text{Cov}(\hat{\beta}) = \sigma^2 (\mathbf{X}^T \mathbf{X})^{-1}$	19
3.8.3	Properties of \hat{y} : $\mathbb{E}(\hat{y}) = \mathbf{X}\beta$, $\text{Cov}(\hat{y}) = \sigma^2 \mathbf{H}$	19
3.8.4	Properties of \mathbf{r} : $\mathbb{E}(\mathbf{r}) = 0$, $\text{Cov}(\mathbf{r}) = \sigma^2 (\mathbf{I}_n - \mathbf{H})$	19
3.8.5	Unbiased Variance Estimation $\mathbb{E}(\hat{\sigma}^2) = \sigma^2$	20
3.9	The Gauss-Markov Theorem: LS estimator is the BLUE (best linear unbiased estimator)	20
3.10	Maximum Likelihood Estimation, Distribution of LS estimates	20
3.10.1	Assumption $\mathbf{y} \sim \mathbf{N}_n (\mathbf{X}\beta, \sigma^2 \mathbf{I})$	20
3.10.2	LS estimator is the Maximum Likelihood Estimator (MLE)	21
3.10.3	Estimations $\hat{\beta} \sim \mathbf{N}_{\mathbf{P}} \left(\beta, \sigma^2 (\mathbf{X}^T \mathbf{X})^{-1} \right)$, $\hat{\mathbf{y}} \sim \mathbf{N}_{\mathbf{n}} (\mathbf{X}\beta, \sigma^2 \mathbf{H})$, $\hat{\mathbf{r}} \sim \mathbf{N}_{\mathbf{n}} (\mathbf{0}, \sigma^2 (\mathbf{I}_n - \mathbf{H}))$	21
3.11	\mathbf{r} Residuals' Properties	21
3.11.1	$n - p$ dimensions: $\mathbf{r} \in \mathbb{R}^{n-p}$	21
3.11.2	Distribution: $\hat{\sigma}^2 \sim \sigma^2 \frac{\chi_{n-p}^2}{n-p}$	21
3.11.3	$\hat{\mathbf{y}}$ and \mathbf{r} are independent	21
3.12	Testing Predictors (Coefficients)	22
3.12.1	Testing a Single Predictor $H_0 : \beta_j = 0$: t -test	22
3.12.2	Review the degree of freedom	22
3.12.3	Testing all predictors: F -test	22
3.12.4	Partial F -test	23
3.13	Permutation Tests (When the normal distribution hypothesis doesn't hold)	24
3.13.1	Procedure	24
3.13.2	Calculation of the p-value: Monte Carlo method	24
3.14	Confidence Intervals for β_j , Confidence Region for β	25
3.14.1	Confidence Intervals for β_j	25
3.14.2	Confidence Region for β	25
3.15	Confidence/Prediction Intervals for New Observations	25
3.15.1	Confidence Interval for $\mu^* = (x^*)^T \beta$	26
3.15.2	Prediction Interval for $y^* = (x^*)^T \beta + e^*$	26
3.15.3	Mahalanobis distance	27
Chapter 4	MLR: unusual observations	29

4.1	1. High leverage points: $h_i \geq \frac{2p}{n}$	29
4.1.1	Leverage Points	29
4.1.2	Properties of the Leverage: $0 < h_i < 1, \sum_i h_i = p$	30
4.1.3	Fitted Values and Leverage: $Var(\hat{y}_i) = \sigma^2 h_i, Var(r_i) = \sigma^2(1 - h_i)$	30
4.1.4	High-leverage Points: $h_i \geq \frac{2p}{n}$	30
4.2	Residuals: Standardized Residuals vs. Studentized residuals	31
4.2.1	Difference between ε and \mathbf{r}	31
4.2.2	Standardized Residuals: $r_i^* = \frac{r_i}{\hat{\sigma}\sqrt{1-h_i}}$	31
4.2.3	Studentized Residuals: $t_i = r_i^* \left(\frac{n-p-1}{n-p-r_i^{*2}} \right)^{1/2}$	31
4.3	2. Outlier (Large $ r_i^* $)	32
4.3.1	Outlier test	32
4.3.2	Bonferroni Correction	32
4.3.3	What we should do with outliers?	32
4.4	3. Highly Influential Points: Large $D_i = \frac{r_i^{*2}}{p} \left(\frac{h_i}{1-h_i} \right)$ ($D_i \geq 1$)	33
4.4.1	Influential observations	33
Chapter 5	MLR Diagnostics: Checking Assumptions	34
5.1	Classical Linear Model (CLM) assumption: Gauss-Markov Assumption + $\varepsilon \sim^{IID} \mathcal{N}(0, \sigma^2)$	34
5.2	Check 1. Constancy of Variance(Homoscedasticity)	34
5.2.1	Method 1: graph <i>residuals</i> against <i>Fitted Values</i> \hat{y}	34
5.2.2	Method 2: Breusch-Pagan Test	35
5.2.3	What happen if Heteroscedasticity	35
5.2.4	What can we do if Heteroscedasticity	36
5.2.5	Remedial measure: Variance Stabilizing Transformations $\sqrt{Y}, \log Y, \frac{1}{Y}$ or $\frac{1}{Y+1}$	36
5.3	Check 2. Normality	37
5.3.1	Method 1: Histogram, graph <i>residuals</i> against its frequency	37
5.3.2	Method 2: QQ-Plot, graph <i>residuals</i> against its frequency	37
5.3.3	Method 3: Shapiro-Wilk Test (good for $n \leq 50$)	38
5.3.4	Method 4: Kolmogorov-Smirnov Test (good for $n > 50$)	38
5.3.5	Remedial measure: Box-Cox Transformations of Y	39
5.4	Checking 3. Serial Dependence	40
5.4.1	Method 1: graph <i>residuals</i> against index variable(time or case number)	40
5.4.2	Method 2: Durbin Watson test	40

5.5	Checking 4. Non-Linearity	41
5.5.1	Method 1: Partial Regression Plots	41
5.5.2	Remedial measure: Linearizing Transformations	41
5.5.3	Remedial measure: Box-Cox Transformations of Y also works	41
Chapter 6	MLR Diagnostics: Collinearity	42
6.1	Exact Collinearity/ linearly dependent	42
6.2	What happens if exact collinearity	42
6.3	Approximate Collinearity	42
6.4	What happens if approximate collinearity: based on $\left(\frac{1}{1-R_k^2}\right)$ (k -th variance inflation factor (VIF), $VIF > 10$)	42
6.5	Possible symptoms of collinearity	43
6.6	Global Measure of Collinearity: <i>condition number</i> of $\mathbf{X}^T \mathbf{X}$	43
6.7	What to do with collinearity	43
Chapter 7	Generalized Least Squares (GLS)	44
7.1	GLS, Σ known ($\hat{\beta} = (\mathbf{X}^\top \Sigma^{-1} \mathbf{X})^{-1} \mathbf{X}^\top \Sigma^{-1} \mathbf{y}$, $RSS = (\mathbf{y} - \mathbf{X}\beta)^\top \Sigma^{-1} (\mathbf{y} - \mathbf{X}\beta)$)	44
7.1.1	Method 1: Transform back to OLS	44
7.1.2	Weighted Least Squares (WLS)	45
7.1.3	WLS special example : Replicated Observations	45
7.1.4	Method 2: Likelihood Estimation	46
7.2	GLS, Σ unknown	46
7.2.1	Method 1: Estimation of Variance (r_i^2)/Standard Deviation Function ($ r_i $)	46
7.2.2	Method 2: iterative approach	47
Chapter 8	GLS Diagnostics: Lack of Fit Tests	48
8.1	Gaussian Assumption	48
8.2	When σ^2 is known	48
8.3	When σ^2 is unknown	49
8.3.1	Hypothesis	49
8.3.2	Under the null hypothesis H_0	49
8.3.3	Under the alternative big-model hypothesis H_α :	49
8.3.4	F -Test	49

Chapter 9 Polynomials Regression	50
9.1 Basic Function	50
9.2 Choose Order d	50
9.3 Orthogonal Polynomials	50
9.4 Piece-wise Polynomials	51
9.5 Cubic Splines	51
9.5.1 Why Spline	51
9.5.2 Settings	51
9.5.3 Number of free parameters: $m + 4$	52
9.5.4 Properties: linear combination also cubic spline	52
9.5.5 Cubic Splines Basis	52
9.6 Natural Cubic Splines (NCS)	53
9.6.1 Degree of Freedom (Number of free parameters): m	53
9.6.2 NCS Basis	53
9.6.3 Note: Waste of Data points	54
9.7 Regression Splines	54
9.8 K -Fold Cross-Validation	54
Chapter 10 Analysis of COVariance (ANOVA): Basic	55
10.1 Two level example	55
10.2 Two-level Test: t-test	56
10.3 Multi-level example	56
10.4 Multi-level Test: F-test	56
Chapter 11 Model Adjustment: Variable Selection	58
11.1 Training and Test Errors	58
11.2 Model selection procedures	59
11.2.1 Testing-based procedures	59
11.2.2 Criterion-based procedures	60
Chapter 12 Model Adjustment: Shrinkage Methods	62
12.1 Principal Components Regression (PCR)	62
12.1.1 Principal Component Analysis (PCA)	62
12.1.2 After PCA	63

12.1.3 Use How many Principal Components?	63
12.2 Ridge Regression	64
12.3 Lasso Regression	65
12.4 Comparing Ridge Regression and Lasso	65
Chapter 13 ANOVA: Comparative Experiments	67
13.1 Terminology	67
13.2 Data	67
13.3 ANOVA Model	67
13.3.1 ANOVA Means Model (Cell Means Model)	67
13.3.2 Factor Effects Model	68
13.4 Model Properties	68
13.5 Model Estimation	69
13.6 F-test	69
13.7 Diagnostics for ANOVA Models	70
13.8 Inference for Factor Level Means (function about the μ_i s)	70
13.8.1 A single factor level mean	70
13.8.2 A difference between two factor level means	71
13.8.3 A contrast among factor level means	71
13.8.4 A linear combination of factor level means	72
13.9 Limitations of Inference Procedures	72
13.10 Bonferroni Correction $\frac{\alpha}{m}$	73
13.11 Tukey's Paired Comparison Procedures	73
13.12 Scheffe's Method for Contrasts	74
Chapter 14 Two Way ANOVA	75
14.1 Factor Effects Model for Two Factors	75
14.2 Interaction Plots	76
14.3 Partitioning of Total Sum of Squares	76
14.4 Partitioning of Treatment Sum of Squares	77
14.5 ANOVA Table	77
14.6 F-test	77
14.7 Estimation of Factor Level Means	78

14.8 Estimation of Treatment Means	79
Chapter 15 Two Way ANOVA: Special Cases	80
15.1 Unbalanced ANOVA (Use Partial F–Test or ANOVA type III in R)	80
15.1.1 Use Partial F–Test	80
15.1.2 ANOVA type III in R	80
15.2 Balanced ANOVA with $n = 1$ (Tukey’s Test)	81
15.2.1 Tukey’s Test for Additivity	81
Chapter 16 Experimental Designs: Introduction	83
16.1 Experimental vs. Observational Study	83
16.2 Principles of Experimental Design	83
16.3 Randomization Test	83
Chapter 17 Experimental Designs: Blocking	84
17.1 Randomized Complete Block Design (RCBD) Model	84
17.2 Latin Squares	85
17.2.1 Example	85
17.2.2 Features of a Latin Square Design	85
17.2.3 Randomization in Latin Square Designs	86
17.2.4 Latin Square Model	86
17.3 Balanced Incomplete Block Design (BIBD)	87
17.4 BIBD Remarks	88
Chapter 18 Linear Models with Random Effects	90
18.1 Random Effects	90
18.2 Intraclass Correlation	90
18.3 Mixed Models	91
18.4 REML: restricted maximum likelihood	91
18.5 ML Estimation	92
18.6 Testing and Confidence Intervals	93
18.7 Testing the Random Effect Variance	93
18.8 Parametric Bootstrap	94

Chapter 1 Review of statistics

1.1 Random Vectors

Suppose a random vector $\vec{Z} \in \mathbb{R}^{m \times 1}$ has mean $\vec{\mu} = \mathbb{E}(\mathbf{Z}) = [\mathbb{E}(Z_1), \mathbb{E}(Z_2), \dots, \mathbb{E}(Z_m)]^T$ and variance-covariance matrix: $\Sigma_{m \times m} = Cov(\mathbf{Z}) = \mathbb{E}((\mathbf{Z} - \mu)(\mathbf{Z} - \mu)^T) = \begin{bmatrix} Var(Z_1) & \cdots & Cov(Z_1, Z_m) \\ \vdots & \ddots & \vdots \\ Cov(Z_m, Z_1) & \cdots & Var(Z_m) \end{bmatrix}$.

1.2 Affine Transformation

(1) The affine transformation $\mathbf{W} = \mathbf{a}_{n \times 1} + \mathbf{B}_{n \times m} \mathbf{Z}_{m \times 1}$ should have following properties:

$$\mathbb{E}(\mathbf{W}) = \mathbf{a} + \mathbf{B}\mu, \quad Cov(\mathbf{W}) = \mathbf{B}\Sigma\mathbf{B}^T$$

(2) The affine transformation $\mathbf{W} = \mathbf{v}^T \mathbf{Z} = v_1 Z_1 + \dots + v_m Z_m$ should have following properties:

$$\begin{aligned} \mathbb{E}(\mathbf{W}) &= \mathbf{v}^T \mu = \sum_{i=1}^m v_i \mu_i \\ Var(\mathbf{W}) &= \mathbf{v}^T \Sigma \mathbf{v} = \sum_{i=1}^m v_i^2 Var(Z_i) + 2 \sum_{i < j} v_i v_j Cov(Z_i, Z_j) \end{aligned}$$

$$\text{i.e. } \mathbb{E}(\mathbf{A}\mathbf{Z}) = \mathbf{A}\mathbb{E}(\mathbf{Z}); \quad Var(\mathbf{A}\mathbf{Z}) = \mathbf{A}Var(\mathbf{Z})\mathbf{A}^T$$

(3)

$$\begin{aligned} Cov(\mathbf{AX}, \mathbf{BY}) &= \mathbb{E}[(\mathbf{AX} - \mathbf{A}\mathbb{E}(X))(\mathbf{BY} - \mathbf{B}\mathbb{E}(Y))^T] \\ &= \mathbf{A}\mathbb{E}[(\mathbf{X} - \mathbb{E}(X))(\mathbf{Y} - \mathbb{E}(Y))^T]\mathbf{B}^T \\ &= \mathbf{ACov}(\mathbf{X}, \mathbf{Y})\mathbf{B}^T \end{aligned}$$

Chapter 2 Regression Analysis (SLR)

2.1 Simple Linear Regression

It is a "tool" used to examine the relationship between **one Dependent Variable or Response** y , and one (or more) **Independent Variables or Regressors or Predictors** X .

$$y = \beta_0 + \beta_1 X$$

where β_0 is the *intercept*; β_1 is the *slope*. One *response* y ; One *predictor* X .

y is a random variable that has a distribution for every level of the independent variable. The data come in pairs:

$$(X, y) = \{(X_i, y_i)\}_{i=1,\dots,n}.$$

Definition 2.1 (Simple Linear Regression Model)

Suppose every data observation can be interpreted as

$$y_i = \beta_0 + \beta_1 X_i + \varepsilon_i$$

where the intercept β_0 , the slope β_1 , and the error term ε_i of observation (X_i, y_i) . (intercept β_0 , the slope β_1 , and error variance σ^2 are the model parameters.)



2.1.1 Assumptions

Definition 2.2 (Assumptions (errors ε form))

The errors $\varepsilon_1, \varepsilon_2, \dots, \varepsilon_n$ are assumed to

- have mean zero: $E(\varepsilon_i) = 0$.
- be uncorrelated: $Cov(\varepsilon_i, \varepsilon_j) = 0, i \neq j$.
- be homoscedastic: $Var(\varepsilon_i) = \sigma^2$ does not depend on i .



The last two could be combined and written as: $Cov(\varepsilon_1, \varepsilon_j) = \sigma^2 \delta_{ij}$ where $\delta_{ij} = \begin{cases} 0 & i \neq j \\ 1 & i = j \end{cases}$

Definition 2.3 (Assumptions ($Y|X$ form))

Based on the SLR model moment assumptions on the error terms, we have the following **equivalent** assumptions for the moments of Y conditioning on X :

- $E(y_i|X_i) = \beta_0 + \beta_1 x_i$
- $Var(y_i|X_i) = \sigma^2$

- $Cov(y_i, y_j | X_i, X_j) = 0, i \neq j$



Claim 2.1 (Interpretation of β_1, β_0)

β_1 is the **change in the mean** of the probability distribution function of y per unit change in x .

β_0 is the **mean** of the probability distribution function of y (at $x = 0$) when $X = \vec{0}$, otherwise β_0 has no particular meaning.



2.1.2 Least Squares Estimations

Suppose the **Residual Sum of Squares (RSS)** is $RSS = \sum_{i=1}^n (y_i - \beta_0 - \beta_1 x_i)^2$. We want to find estimates of β_0, β_1 to minimize the RSS:

$$(\hat{\beta}_0, \hat{\beta}_1) = \underset{(\beta_0, \beta_1)}{\operatorname{argmin}} RSS$$

Proposition 2.1 (LS Estimators)

$$\hat{\beta}_0 = \bar{y} - \hat{\beta}_1 \bar{x};$$

$$\hat{\beta}_1 = \frac{\sum_{i=1}^n x_i y_i - n \bar{x} \bar{y}}{\sum_{i=1}^n x_i^2 - n \bar{x}^2} = \frac{\sum_{i=1}^n (x_i - \bar{x})(y_i - \bar{y})}{\sum_{i=1}^n (x_i - \bar{x})^2} = \frac{\sum_{i=1}^n (x_i - \bar{x}) y_i}{\sum_{i=1}^n (x_i - \bar{x})^2}$$

Alternative Representation of $\hat{\beta}_1$ is

$$\hat{\beta}_1 = \frac{\sum_i (x_i - \bar{x})(y_i - \bar{y})}{\sum_i (x_i - \bar{x})^2} = \frac{S_{xy}}{S_{xx}} = r_{xy} \sqrt{\frac{S_{yy}}{S_{xx}}}$$

where $S_{xy} = \sum_i (x_i - \bar{x})(y_i - \bar{y})$, $S_{xx} = \sum_i (x_i - \bar{x})^2$, $S_{yy} = \sum_i (y_i - \bar{y})^2$, and $r_{xy} = \frac{\sum_i (x_i - \bar{x})(y_i - \bar{y})}{\sqrt{\sum_i (x_i - \bar{x})^2} \sqrt{\sum_i (y_i - \bar{y})^2}} = \frac{S_{xy}}{\sqrt{S_{xx} S_{yy}}}$.



Proof 2.1

$$\frac{\partial RSS}{\partial \beta_0} = 0 \Leftrightarrow -2 \sum_{i=1}^n (y_i - \beta_0 - \beta_1 x_i) = 0 \Leftrightarrow \beta_0 n + \beta_1 \sum_{i=1}^n x_i = \sum_{i=1}^n y_i$$

$$\frac{\partial RSS}{\partial \beta_1} = 0 \Leftrightarrow -2 \sum_{i=1}^n (y_i - \beta_0 - \beta_1 x_i) x_i = 0 \Leftrightarrow \beta_0 \sum_{i=1}^n x_i + \beta_1 \sum_{i=1}^n x_i^2 = \sum_{i=1}^n x_i y_i$$

Lemma 2.1 (Unbiased Estimators)

$$\mathbb{E}\hat{\beta}_0 = \beta_0, \mathbb{E}\hat{\beta}_1 = \beta_1.$$



Proof 2.2

$$\mathbb{E}\hat{\beta}_1 = \mathbb{E} \left[\frac{\sum_{i=1}^n (x_i - \bar{x})(\beta_0 + \beta_1 x_i + \varepsilon_i)}{\sum_{i=1}^n (x_i - \bar{x})^2} \right] = \mathbb{E} \left[\frac{\sum_{i=1}^n (x_i - \bar{x}) \beta_1 x_i}{\sum_{i=1}^n (x_i - \bar{x})^2} \right] = \beta_1. \quad \mathbb{E}\hat{\beta}_0 = \mathbb{E} [\bar{y} - \hat{\beta}_1 \bar{x}] = \beta_0.$$

2.1.3 Fitted Values & Residuals

Definition 2.4 (Fitted Values)

The Prediction of y_i or the fitted value at x_i is

$$\hat{y}_i = \hat{\beta}_0 + \hat{\beta}_1 x_i$$



$$\begin{aligned}\hat{y}_i &= \hat{\beta}_0 + \hat{\beta}_1 x_i \\ &= \bar{y} - \hat{\beta}_1 \bar{x} + \hat{\beta}_1 x_i \\ &= \bar{y} + \frac{\sum_{i=1}^n (x_i - \bar{x})(y_i - \bar{y})}{\sum_{i=1}^n (x_i - \bar{x})^2} (x_i - \bar{x})\end{aligned}$$

Definition 2.5

The i^{th} residual is $r_i = y_i - \hat{y}_i$.



Proposition 2.2 (Properties of residuals)

1. $\sum_i r_i = 0$.
2. $RSS = \sum_i r_i^2$ is minimized.
3. $\sum_{i=1}^n y_i = \sum_{i=1}^n \hat{y}_i$.
4. $\sum_i x_i r_i = 0$ prove by first-order condition (proof: $\sum_i x_i r_i = \sum_i x_i (y_i - \bar{y} - \hat{\beta}_1(x_i - \bar{x})) = \sum_i x_i y_i - n\bar{x}\bar{y} - \frac{\sum_{i=1}^n x_i y_i - n\bar{x}\bar{y}}{\sum_{i=1}^n x_i^2 - n\bar{x}^2} (\sum_i x_i^2 - n\bar{x}^2) = 0$) (We can interpret Least Square estimates Y 's values in X 's direction, the residual should be perpendicular to X .)
5. $\sum_i \hat{y}_i r_i = 0$ (inferred from 4).
6. The regression line always goes through the point (\bar{x}, \bar{y}) .



2.1.4 Degree of freedom

The **degree of freedom(df)** of the residuals is

$$df = (\text{Sample size}) - (\# \text{ of parameters})$$

$df = n - 2$ in this case.

2.1.5 (Sample) Error variance

The error variance is estimated by

$$\hat{\sigma}^2 = \frac{1}{n-2} \sum_i r_i^2$$

2.2 Goodness of Fit: R -square

2.2.1 TSS, RSS, FSS

$$TSS : \sum_i (y_i - \bar{y})^2$$

$$RSS : \sum_i r_i^2$$

$$FSS : \sum_i (\hat{y}_i - \bar{y})^2$$

$$\begin{aligned} \sum_i (y_i - \bar{y})^2 &= \sum_i (y_i - \hat{y}_i + \hat{y}_i - \bar{y})^2 = \sum_i (r_i + \hat{y}_i - \bar{y})^2 \\ &= \sum_i r_i^2 + \sum_i (\hat{y}_i - \bar{y})^2 \end{aligned}$$

$$TSS = RSS + FSS$$

2.2.2 Coefficient of Determination(R^2)

$$R^2 = \frac{\sum_i (\hat{y}_i - \bar{y})^2}{\sum_i (y_i - \bar{y})^2} = \frac{FSS}{TSS} = \frac{TSS - RSS}{TSS} = 1 - \frac{RSS}{TSS}$$

$$0 \leq R^2 \leq 1$$

It measures the effect of X in reducing the variation in Y .

The larger R^2 is, the more the total variation of y is reduced by reducing the independent variable x .

R^2 can also represent the degree of linear association between X and Y .

$r_{xy} = \pm \sqrt{R^2}$, where the sign is the sign of the slope.

$$\begin{aligned} r_{xy}^2 &= \frac{(\sum_i (x_i - \bar{x})(y_i - \bar{y}))^2}{\sum_i (x_i - \bar{x})^2 \sum_i (y_i - \bar{y})^2} = \frac{(\sum_i (x_i - \bar{x})(r_i + \hat{y}_i - \bar{y}))^2}{\sum_i (x_i - \bar{x})^2 \sum_i (y_i - \bar{y})^2} \\ &= \frac{(\sum_i (x_i - \bar{x})(\hat{y}_i - \bar{y}))^2}{\sum_i (x_i - \bar{x})^2 \sum_i (y_i - \bar{y})^2} = \frac{(\sum_i (\hat{\beta}_1 x_i - \hat{\beta}_1 \bar{x})(\hat{y}_i - \bar{y}))^2}{\sum_i (\hat{\beta}_1 x_i - \hat{\beta}_1 \bar{x})^2 \sum_i (y_i - \bar{y})^2} \\ &= \frac{(\sum_i (\hat{y}_i - \bar{y}))^2}{\sum_i (\hat{y}_i - \bar{y})^2 \sum_i (y_i - \bar{y})^2} = \frac{\sum_i (\hat{y}_i - \bar{y})^2}{\sum_i (y_i - \bar{y})^2} = R^2 \end{aligned}$$

2.3 Affine Transformations

Suppose we have a SLR model of Y on X , i.e. $y_i = \beta_0 + \beta_1 x_i$

2.3.1 $\tilde{y}_i = ay_i + b$

1. Rescale y_i by $\tilde{y}_i = ay_i + b$ and then regress \tilde{y}_i on x_i . How would the LS estimates and R^2 be affected?

$$\begin{aligned}\tilde{\beta}_1 &= \frac{\sum_{i=1}^n (x_i - \bar{x})(ay_i + b - a\bar{y} - b)}{\sum_{i=1}^n (x_i - \bar{x})^2} = a\hat{\beta}_1 \\ \tilde{\beta}_0 &= a\bar{y} + b - \tilde{\beta}_1\bar{x} = a\hat{\beta}_0 + b \\ \tilde{R}^2 &= \frac{\sum_i (a\hat{y}_i + b - a\bar{y} - b)^2}{\sum_i (ay_i + b - a\bar{y} - b)^2} = R^2\end{aligned}$$

2.3.2 $\tilde{x}_i = ax_i + b$

2. Rescale y_i by $\tilde{x}_i = ax_i + b$ and then regress y_i on \tilde{x}_i . How would the LS estimates and R^2 be affected?

$$\begin{aligned}\tilde{\beta}_1 &= \frac{\sum_{i=1}^n (ax_i + b - a\bar{x} - b)(y_i - \bar{y})}{\sum_{i=1}^n (ax_i + b - a\bar{x} - b)^2} = \frac{\hat{\beta}_1}{a} \\ \tilde{\beta}_0 &= \bar{y} - \tilde{\beta}_1(a\bar{x} + b) = \hat{\beta}_0 - \frac{b}{a}\hat{\beta}_1 \\ \tilde{R}^2 &= \frac{\sum_i (\hat{y}_i - \bar{y})^2}{\sum_i (y_i - \bar{y})^2} = R^2\end{aligned}$$

2.3.3 Regress x on y instead

3. Regress x on y instead

$$x = \tilde{\beta}_0 + \tilde{\beta}_1 y$$

$$\tilde{\beta}_1 = \frac{S_{xy}}{S_{yy}}; \quad \tilde{\beta}_0 = \bar{x} - \tilde{\beta}_1 \bar{y}; \quad \tilde{R}^2 = r_{xy}^2 = R^2$$

2.4 Regression Through the Origin

$$y_i \approx \beta_1 x_i$$

(1) $\hat{\beta}_1$:

By LS: $\min_{\hat{\beta}_1} RSS = \sum_i (\hat{\beta}_1 x_i - y_i)^2$

$$\frac{\partial RSS}{\partial \hat{\beta}_1} = \sum_i 2x_i(\hat{\beta}_1 x_i - y_i) = 0 \Rightarrow \hat{\beta}_1 = \frac{\sum_{i=1}^n x_i y_i}{\sum_{i=1}^n x_i^2}$$

(2) R^2 :

negative R^2 is possible since $R^2 = 1 - \frac{RSS}{TSS}$ and RSS may be larger than TSS .

We use a modified R-square

$$\sum_i y_i^2 = \sum_i (y_i - \hat{y}_i + \hat{y}_i)^2 = \sum_i (y_i - \hat{y}_i)^2 + \sum_i \hat{y}_i^2$$

$$\tilde{R}^2 = \frac{\sum_i \hat{y}_i^2}{\sum_i y_i^2} = 1 - \frac{\sum_i (y_i - \hat{y}_i)^2}{\sum_i y_i^2} = 1 - \frac{RSS}{\sum_i y_i^2}$$

2.5 LS Estimators Properties

2.5.1 Unbiasedness of LS Estimators $E(\hat{\beta}_1) = \beta_1, E(\hat{\beta}_0) = \beta_0$

x_i 's (\mathcal{X}) are already known.

$$\begin{aligned}\mathbb{E}_y(\hat{\beta}_1) &= \mathbb{E}\left[\frac{\sum_i (x_i - \bar{x}) y_i}{\sum_i (x_i - \bar{x})^2}\right] = \frac{\sum_i (x_i - \bar{x}) \cdot \mathbb{E}(y_i)}{\sum_i (x_i - \bar{x})^2} \\ &= \frac{\sum_i (x_i - \bar{x}) \cdot \mathbb{E}(\beta_0 + \beta_1 x_i)}{\sum_i (x_i - \bar{x})^2} = \sum_i c_i (\beta_0 + \beta_1 x_i) \\ &= \beta_0 \sum c_i + \beta_1 \sum c_i x_i = \beta_1, \text{ where } c_i = \frac{(x_i - \bar{x})}{\sum_i (x_i - \bar{x})^2} \\ \mathbb{E}(\hat{\beta}_0) &= \mathbb{E}(\bar{y} - \hat{\beta}_1 \bar{x}) \\ &= \mathbb{E}(\bar{y}) - \bar{x} \cdot \mathbb{E}(\hat{\beta}_1) = \frac{1}{n} \sum_i \mathbb{E}(y_i) - \bar{x} \cdot \beta_1 \\ &= \frac{1}{n} \sum_i \mathbb{E}(\beta_0 + \beta_1 x_i) - \bar{x} \cdot \beta_1 \\ &= \beta_0 + \bar{x} \cdot \beta_1 - \bar{x} \cdot \beta_1 = \beta_0\end{aligned}$$

2.5.2 Mean squared error(MSE) of LS Estimators = $Var(\hat{\beta}_1) = \sigma^2 \frac{1}{S_{xx}}$,

$$Var(\hat{\beta}_0) = \sigma^2 \left(\frac{1}{n} + \frac{\bar{x}^2}{S_{xx}} \right)$$

Mean squared error(MSE)= $\frac{1}{n} \sum_{i=1}^n (y_i - \hat{y}_i)^2$

Note that since both estimators are unbiased \Rightarrow MSE = Variance.

1. MSE for slope

$$\begin{aligned}\text{Var}(\hat{\beta}_1) &= \text{Var}\left[\frac{\sum_i (x_i - \bar{x}) y_i}{\sum_i (x_i - \bar{x})^2}\right] = \text{Var}\left(\sum_i c_i y_i\right) (c_i \text{ as before }) \\ &= \sum_i c_i^2 \cdot \text{Var}(y_i) = \sum_i c_i^2 \sigma^2 (\text{ from model assumption }) \\ &= \sigma^2 \cdot \left(\frac{\sum_i (x_i - \bar{x})}{\sum_i (x_i - \bar{x})^2}\right)^2 = \frac{\sigma^2}{\sum_i (x_i - \bar{x})^2} = \sigma^2 \frac{1}{S_{xx}}\end{aligned}$$

2. MSE for intercept

$$\text{Var}(\hat{\beta}_0) = \text{Var}(\bar{y} - \hat{\beta}_1 \bar{x}) = \sigma^2 \left(\frac{1}{n} + \frac{\bar{x}^2}{S_{xx}} \right)$$

2.6 Normality Assumption

Additionally, we assume that

$$\varepsilon_i \sim^{iid} \mathcal{N}(0, \sigma^2)$$

Equivalently,

$$y_i \sim^{iid} \mathcal{N}(\beta_0 + \beta_1 x_i, \sigma^2)$$

(y_i are a linear shift of the ε_i , so it is also normally distributed)

(The y_i 's are jointly normal, and so are linear combinations of the y_i 's, since the errors are normally distributed and uncorrelated/independent.)

2.7 Distribution of LS Estimators

$\hat{\beta}_1$ and $\hat{\beta}_0$ are jointly normally distributed with

$$\begin{aligned}\mathbb{E}(\hat{\beta}_1) &= \beta_1 & \text{Var}(\hat{\beta}_1) &= \sigma^2 \frac{1}{S_{xx}} \\ \mathbb{E}(\hat{\beta}_0) &= \beta_0 & \text{Var}(\hat{\beta}_0) &= \sigma^2 \left(\frac{1}{n} + \frac{\bar{x}^2}{S_{xx}} \right) \\ \text{Cov}(\hat{\beta}_1, \hat{\beta}_0) &= -\sigma^2 \frac{\bar{x}}{S_{xx}}\end{aligned}$$

$RSS = \sum_i (y_i - \hat{y}_i)^2 \sim \sigma^2 \chi_{n-2}^2$ which implies that

$$\mathbb{E}(\hat{\sigma}^2) = \mathbb{E}\left(\frac{RSS}{n-2}\right) = \frac{\sigma^2(n-2)}{n-2} = \sigma^2$$

$(\hat{\beta}_0, \hat{\beta}_1)$ and RSS are independent.

2.8 Hypothesis Testing (T-test)

2.8.1 Testing for the Slope

$$\begin{cases} H_0 : \beta_1 = c(\text{null}) \\ H_\alpha : \beta_1 \neq c (\text{alternative}) \end{cases}$$

where c is an known constant. The test statistics is

$$t = \frac{\hat{\beta}_1 - c}{\sqrt{\text{Var}(\hat{\beta}_1)}} = \frac{\hat{\beta}_1 - c}{\hat{\sigma}/\sqrt{S_{xx}}}$$

The distribution of t under the null is T_{n-2} . The p -value is twice the area under the T_{n-2} distribution more extreme than the observed statistic t .

2.8.2 Testing for the Intercept

$$\begin{cases} H_0 : \beta_0 = c(\text{null}) \\ H_\alpha : \beta_0 \neq c (\text{alternative}) \end{cases}$$

The test statistics is

$$t = \frac{\hat{\beta}_0 - c}{\sqrt{\text{Var}(\hat{\beta}_0)}}$$

The distribution of t under the null is T_{n-2} . The p -value is twice the area under the T_{n-2} distribution more extreme than the observed statistic t .

2.9 ANOVA Table & F-Test

2.9.1 Degrees of Freedom

$df_{TSS} = n - 1$: one df is lost, because the sample mean is used to estimate the population mean.

$df_{RSS} = n - 2$: two df are lost, because the two parameters are estimated in obtaining the fitted values \hat{y}

$df_{FSS} = 1$: there are n deviations $\hat{y}_i - \bar{y}$, but all the fitted values are associated with the same regression line.

$$df_{TSS} = df_{RSS} + df_{FSS}$$

Sum of Squares	Expression	df
TSS	$\sum_i (y_i - \bar{y})^2$	$n - 1$
FSS	$\sum_i (\hat{y}_i - \bar{y})^2$	1
RSS	$\sum_i (y_i - \hat{y}_i)^2$	$n - 2$

2.9.2 ANOVA Table

Source	SS	df	MS	F
Regression (model)	FSS	1	$MSReg = \frac{FSS}{1}$	$F = \frac{MSReg}{MSE}$
Error	RSS	$n - 2$	$MSE = \frac{RSS}{n-2}$	
Total	TSS	$n - 1$		

Figure 2.1

2.9.3 F-Test (equivalent to t-test)

An alternative way to test for the model parameters is using the F test:

$$\begin{cases} H_0 : \beta_1 = 0 \\ H_\alpha : \beta_1 \neq 0 \end{cases}$$

- Under H_0 , the F -test statistic is

$$F = \frac{MSReg}{MSE} = \frac{FSS}{RSS/(n-2)} \sim F_{1,n-2}$$

- It can be shown that the F -test statistic is equal to the square of the t -test statistic and their p -values are the same. So, **this test is equivalent to the t -test before.**

2.10 Estimation and Prediction

2.10.1 Estimation (always reported for a parameter: $\beta_0 + \beta_1 x^* = \mathbb{E}(Y|x^*)$)

1. Estimation: We want to estimate the mean response at x^* . This is equivalent to estimate: $\beta_0 + \beta_1 x^*$
2. Accuracy of the estimation: is measured by the expected value of the squared difference between the point estimate and the target.

- For estimation the target is $\beta_0 + \beta_1 x^*$:

$$\begin{aligned} & \mathbb{E} (\hat{\beta}_0 + \hat{\beta}_1 x^* - \beta_0 - \beta_1 x^*)^2 \\ &= \text{Var} (\hat{\beta}_0 + \hat{\beta}_1 x^*) \\ &= \text{Var} (\hat{\beta}_0) + (x^*)^2 \text{Var} (\hat{\beta}_1) + 2x^* \text{Cov} (\hat{\beta}_0, \hat{\beta}_1) \\ &= \sigma^2 \left(\frac{1}{n} + \frac{(x^* - \bar{x})^2}{\sum_i (x_i - \bar{x})^2} \right) \\ &= \sigma^2 \left(\frac{1}{n} + \frac{(x^* - \bar{x})^2}{S_{xx}} \right) \end{aligned}$$

3. Confidence interval: An $(1 - \alpha)100\%$ Confidence Interval for the Mean Response when $x = x^*$ is given by

$$\hat{\beta}_0 + \hat{\beta}_1 x^* \pm T_{n-2}(\alpha/2) \hat{\sigma} \sqrt{\frac{1}{n} + \frac{(x^* - \bar{x})^2}{S_{xx}}}$$

2.10.2 Prediction (is reported for the value of a random variable Y^*)

1. Prediction: of an outcome of random variable Y^* at a given value x^* , where $Y^* \sim N(\beta_0 + \beta_1 x^*, \sigma^2)$
2. For prediction the target is $Y^* = \beta_0 + \beta_1 x^* + e^*$, where $e^* \sim N(0, \sigma^2)$ This new error e^* is independent of the previous n data points, i.e. is independent of $(\hat{\beta}_0, \hat{\beta}_1)$

$$\begin{aligned} & \mathbb{E} \left[(\hat{\beta}_0 + \hat{\beta}_1 x^* - Y^*)^2 \right] \\ &= \mathbb{E} \left[(\hat{\beta}_0 + \hat{\beta}_1 x^* - \beta_0 - \beta_1 x^* - e^*)^2 \right] \\ &= \mathbb{E} \left[(\hat{\beta}_0 + \hat{\beta}_1 x^* - \beta_0 - \beta_1 x^*)^2 \right] + \mathbb{E} [(e^*)^2] \\ &= \sigma^2 \left(1 + \frac{1}{n} + \frac{(x^* - \bar{x})^2}{S_{xx}} \right) \end{aligned}$$

3. Prediction interval: An $(1 - \alpha)100\%$ Prediction Interval for \hat{Y}^* when $x = x^*$ is given by

$$\hat{\beta}_0 + \hat{\beta}_1 x^* \pm T_{n-2}(\alpha/2) \hat{\sigma} \sqrt{1 + \frac{1}{n} + \frac{(x^* - \bar{x})^2}{S_{xx}}}$$

2.10.3 Simultaneous Confidence

$\mu^* = \mathbb{E}[y|x^*] = \beta_0 + \beta_1 x^*$'s Confidence Interval:

$$I(x^*) = (\hat{\mu}^* \pm T_{n-2}(\frac{\alpha}{2}) se(\hat{\mu}^*))$$

Where

$$\hat{\mu}^* = \hat{\beta}_0 + \hat{\beta}_1 x^* \text{ and } se(\hat{\mu}^*) = \hat{\sigma} \sqrt{\frac{1}{n} + \frac{(x^* - \bar{x})^2}{\sum_{i=1}^n (x_i - \bar{x})^2}}$$

If we want confidence intervals at multiple points $(x_1^*, x_2^*, \dots, x_m^*)$, we can use formula (1) to have confidence intervals at the m points: $I(x_1^*), I(x_2^*), \dots, I(x_m^*)$.

We know that:

$$\mathbb{P}(\mu_i^* \in I(x_i^*)) = (1 - \alpha)$$

This is the point-wise coverage probability for μ_i^* and formula (1) gives the point-wise CI.

What about the simultaneous coverage probability? i.e.:

$$\mathbb{P}(\mu_i^* \in I(x_i^*), \text{ for } i = 1, \dots, m) = ?$$

To make sure that (for example):

$$\mathbb{P}(\mu_i^* \in I(x_i^*), \text{ for } i = 1, \dots, m) = .95$$

we need to set $\alpha = 5\% / m$, which is known as the **Bonferroni correction**.

Let A_k denotes the event that the k th confidence interval covers μ_k^* with:

$$\mathbb{P}(A_k) = (1 - \alpha)$$

Then we can show:

$$\begin{aligned} & \mathbb{P}(\text{ All Cls cover the corresponding } \mu_k^* \text{ values }) \\ &= \mathbb{P}(A_1 \cap A_2 \dots \cap A_m) \\ &= 1 - \mathbb{P}(A_1^c \cup A_2^c \dots \cup A_m^c) \\ &\geq 1 - \mathbb{P}(A_1^c) - \dots - \mathbb{P}(A_m^c) \\ &= 1 - m\alpha \end{aligned}$$

If we choose α/m instead of α , the simultaneous coverage probability will be $(1 - \alpha)$

2.10.4 Confidence Band (Larger than CI)

Ideally we would like to construct a simultaneous confidence band (i.e., $m = \infty$) across all x^* 's. (Scheffé's Theorem - 2959). Let

$$I(x) = (\hat{r}(x) - c\hat{\sigma}, \hat{r}(x) + c\hat{\sigma})$$

where

$$\hat{r}(x) = \hat{\beta}_0 + \hat{\beta}_1 x, c\hat{\sigma} = \sqrt{2F(\alpha, 2, n-2)} \sqrt{\frac{1}{n} + \frac{(x - \bar{x})^2}{\sum_{i=1}^n (x_i - \bar{x})^2}}$$

Then,

$$\mathbb{P}(r(x) \in I(x) \text{ for all } x) \geq 1 - \alpha$$

Can we construct a simultaneous prediction band? No!

Are confidence bands always wider than point-wise confidence intervals? Yes! For SLR, at a location x^* , we have

$$\text{band : } \hat{\mu}^* \pm \sqrt{2F(\alpha, 2, n-2)} \text{ se}(\hat{\mu}^*)$$

$$\text{interval : } \hat{\mu}^* \pm T_{n-2}(\alpha/2) \text{ se}(\hat{\mu}^*)$$

$$\sqrt{2F(\alpha, 2, n-2)} > T_{n-2}(\alpha/2) = \sqrt{2F(\alpha, 1, n-2)}$$

In fact, for any α , we have

$$T_m(\alpha/2) = \sqrt{2F(\alpha, 1, m)} < \sqrt{kF(\alpha, k, m)}$$

2.11 Maximum likelihood estimators with normal error terms

We start with the statistical model, which is the Gaussian-noise simple linear regression model, defined as follows:

1. The distribution of X is arbitrary (and perhaps X is even non-random).
2. If $X = x$, then $Y = \beta_0 + \beta_1 x + \epsilon$, for some constants ("coefficients", "parameters") β_0 and β_1 , and some random noise variable ϵ .
3. $\epsilon \sim N(0, \sigma^2)$, and is independent of X .
4. ϵ is independent across observations.

$$p(y_i|x_i; \beta_0, \beta_1, \sigma^2) = \frac{1}{\sqrt{2\pi\sigma^2}} e^{-\frac{1}{2}(\frac{y_i-(\beta_0+\beta_1x_i)}{\sigma})^2}$$

Given any data set $(x_1, y_1), (x_2, y_2), \dots, (x_n, y_n)$, we can now write down the probability density, under the model, of seeing that data:

$$\prod_{i=1}^n p(y_i|x_i; \beta_0, \beta_1, \sigma^2) = \prod_{i=1}^n \frac{1}{\sqrt{2\pi\sigma^2}} e^{-\frac{1}{2}(\frac{y_i-(\beta_0+\beta_1x_i)}{\sigma})^2} = \frac{1}{(2\pi\sigma^2)^{\frac{n}{2}}} e^{-\frac{\sum_{i=1}^n (y_i-(\beta_0+\beta_1x_i))^2}{2\sigma^2}}$$

Take the **log-likelihood**

$$L(\beta_0, \beta_1, \sigma^2) = -\frac{n}{2} \log 2\pi - n \log \sigma - \frac{1}{2\sigma^2} \sum_{i=1}^n (y_i - (\beta_0 + \beta_1 x_i))^2$$

Then we can compute the **Maximum likelihood estimators** ($\hat{\beta}_0, \hat{\beta}_1, \hat{\sigma}^2$):

(1) $(\hat{\beta}_0, \hat{\beta}_1)$,

Obviously, maximizing $L(\hat{\beta}_0, \hat{\beta}_1, \hat{\sigma}^2)$ is as same as minimizing $\sum_{i=1}^n (y_i - (\hat{\beta}_0 + \hat{\beta}_1 x_i))^2$, then the **Maximum likelihood estimators** is exactly the **LS estimators**:

$$\begin{aligned}\hat{\beta}_1 &= \frac{\sum_{i=1}^n (x_i - \bar{x})(y_i - \bar{y})}{\sum_{i=1}^n (x_i - \bar{x})^2} \\ \hat{\beta}_0 &= \bar{y} - \hat{\beta}_1 \bar{x}\end{aligned}$$

(2) $\hat{\sigma}^2$,

And the $\hat{\sigma}^2$ is exactly the in-sample mean squared error:

$$\hat{\sigma}^2 = \frac{1}{n} \sum_{i=1}^n (y_i - (\hat{\beta}_0 + \hat{\beta}_1 x_i))^2$$

Chapter 3 Multiple Linear Regression (MLR): Basic

3.1 Basic

x_1, x_2, \dots, x_p be p predictors of a response y .

$$y_1 \quad x_{11} \quad x_{12} \quad \cdots \quad x_{1p}$$

$$\begin{array}{cccccc} & y_2 & x_{21} & x_{22} & \cdots & x_{2p} \\ \text{The data will be of the form:} & \vdots & \vdots & \vdots & \ddots & \vdots \end{array}$$

$$y_n \quad x_{n1} \quad x_{n2} \quad \cdots \quad x_{np}$$

Model Equation:

$$y_i = \beta_1 x_{i1} + \beta_2 x_{i2} + \cdots + \beta_p x_{ip} + \varepsilon_i, \quad i = 1, \dots, n$$

where we denote $\mathbf{x}_i = (x_{i1}, \dots, x_{ip})^T$, with $x_{i1} = 1$

$(\beta_1, \beta_2, \dots, \beta_p; \sigma^2)$ are unknown true parameters.

β_1 is the intercept.

$\beta_2, \beta_3, \dots, \beta_p$ are partial slopes.

σ^2 is the error variance

3.1.1 Assumptions of errors

$\varepsilon_1, \varepsilon_2, \dots, \varepsilon_n$ are the random errors. They usually assumed to satisfy the same conditions as in simple linear regression:

- zero mean: $\mathbb{E}(\varepsilon_i) = 0$
- uncorrelated: $\text{Cov}(\varepsilon_i, \varepsilon_j) = 0, i \neq j$, and
- homoscedastic: $\text{Var}(\varepsilon_i) = \sigma^2$ does not depend on i .

3.1.2 Matrix Representation

Matrix Representation of the MLR Model:

$$\begin{array}{ccccccc} \mathbf{y}_{n \times 1} & = & \mathbf{X}_{n \times p} & \beta_{p \times 1} & + & \varepsilon_{n \times 1} \\ \uparrow & & \uparrow & \uparrow & & \uparrow \\ \text{Response} & & \text{Design} & \text{Coefficients} & & \text{Error} \\ & & \text{Matrix} & & & \text{Term} \end{array}$$

- n : sample size
- p : number of predictors or columns of X
- By default the intercept is included in the model in which case the first column of X is a vector of 1's.

We set $\mathbb{E}(\varepsilon) = 0$ and $Cov(\varepsilon) = \sigma^2 \mathbf{I}_n$, then we can infer that

$$\mathbb{E}(y) = \mathbf{X}\beta, \quad Cov(y) = \sigma^2 \mathbf{I}_n$$

3.2 Parameter Estimation $\hat{\beta}$

3.2.1 Parameter Estimation $\hat{\beta} = (\mathbf{X}^T \mathbf{X})^{-1} \mathbf{X}^T y$

- We want to estimate β , i.e. obtain:

$$\hat{\beta} = (\hat{\beta}_1, \hat{\beta}_2, \dots, \hat{\beta}_p)^T$$

- The LS estimator of β minimizes the sum of squared residuals:

$$RSS = \|y - \mathbf{X}\beta\|^2 = (y - \mathbf{X}\beta)^T (y - \mathbf{X}\beta)$$

In order to minimize $RSS = (y - \mathbf{X}\beta)^T (y - \mathbf{X}\beta)$, we take derivatives with respect to β 's and set to zero (as before).

$$\frac{\partial RSS}{\partial \beta} = -2\mathbf{X}_{p \times n}^T (y - \mathbf{X}\beta)_{n \times 1} = \mathbf{0}_{p \times 1}$$

$$\mathbf{X}^T (y - \mathbf{X}\beta) = \mathbf{0} \longrightarrow \text{Normal Equations}$$

$$(\mathbf{X}^T \mathbf{X}) \beta = \mathbf{X}^T y$$

$$\hat{\beta} = (\mathbf{X}^T \mathbf{X})^{-1} \mathbf{X}^T y \rightarrow \text{LS Estimators}$$

Remarks

1. We assume that the rank of X is p , i.e. no columns of X is a linear combinations of the other columns of X .
2. Since \mathbf{X} has rank p , the inverse of $(\mathbf{X}^T \mathbf{X})$ exists.
3. if $(\mathbf{X}^T \mathbf{X})$ is singular the LS solutions is not unique (identifiability problem)

3.2.2 Hat Matrix $\mathbf{H}_{n \times n} = \mathbf{X} (\mathbf{X}^T \mathbf{X})^{-1} \mathbf{X}^T$

Fitted Values $\hat{y}_{n \times 1} = \mathbf{H}_{n \times n} y_{n \times 1}$

$$\begin{aligned} \hat{y}_{n \times 1} &= \mathbf{X} \hat{\beta} \\ &= \mathbf{X} (\mathbf{X}^T \mathbf{X})^{-1} \mathbf{X}^T y \\ &= \mathbf{X} (\mathbf{X}^T \mathbf{X})^{-1} \mathbf{X}^T y \\ &= \mathbf{H}_{n \times n} y_{n \times 1} \end{aligned}$$

$\mathbf{H}_{n \times n} = \mathbf{X} (\mathbf{X}^T \mathbf{X})^{-1} \mathbf{X}^T$ is called the hat matrix, since it returns the "y-hat" values.

3.3 Residuals

3.3.1 Residuals: $\mathbf{r} = (\mathbf{I} - \mathbf{H})\mathbf{y}$

Residuals $\mathbf{r} = (\mathbf{I} - \mathbf{H})\mathbf{y}$, (Sample) error variance $\hat{\sigma}^2 = \frac{\mathbf{r}^T \mathbf{r}}{n-p}$

$$\mathbf{r}_{n \times 1} = \mathbf{y} - \hat{\mathbf{y}} = \mathbf{y} - \mathbf{H}\mathbf{y} = (\mathbf{I} - \mathbf{H})\mathbf{y}$$

3.3.2 Error Variance: $\hat{\sigma}^2 = \frac{\mathbf{r}^T \mathbf{r}}{n-p}$

The residuals \mathbf{r} are used to estimate the error variance:

$$\hat{\sigma}^2 = \frac{1}{n-p} \sum_i r_i^2 = \frac{\mathbf{r}^T \mathbf{r}}{n-p} = \frac{RSS}{n-p}$$

3.4 Properties of Residuals

The LS estimator is the β that satisfies the normal equations, that is

$$\mathbf{X}^T(\mathbf{y} - \hat{\mathbf{y}}) = \mathbf{X}^T(\mathbf{y} - \mathbf{X}\hat{\beta}) = \mathbf{0}$$

This implies the following properties for the residuals, $r_{n \times 1} = \mathbf{y} - \mathbf{X}\hat{\beta}$:

3.4.1 Orthogonal to column of data $\mathbf{X}^T \setminus = 0$

1. The cross-products between the residual vector r and each column of \mathbf{X} are zero, i.e.

$$\begin{aligned} \mathbf{X}^T \mathbf{r} &= \mathbf{X}^T(\mathbf{y} - \mathbf{X}\hat{\beta}) = \mathbf{X}^T \mathbf{y} - \mathbf{X}^T \mathbf{X} \hat{\beta} \\ &= \mathbf{X}^T \mathbf{y} - (\mathbf{X}^T \mathbf{X}) (\mathbf{X}^T \mathbf{X})^{-1} \mathbf{X}^T \mathbf{y} = 0 \end{aligned}$$

3.4.2 Orthogonal to fitted value $\hat{\mathbf{y}}^T \setminus = 0$

2. The cross-product between the fitted value $\hat{\mathbf{y}}$ and the residual vector r is zero, i.e.

$$\hat{\mathbf{y}}^T r = \hat{\beta}^T X^T r = 0$$

This implies that the residual vector r is **orthogonal** to each column of X and $\hat{\mathbf{y}}$.

3.5 Properties of \mathbf{H}

3.5.1 $\mathbf{H}\mathbf{X} = \mathbf{X}$

Let c be any linear combination of the columns of \mathbb{X} , then

$$\mathbf{H}c = c$$

since $\mathbf{H}\mathbf{X} = \mathbf{X} (\mathbf{X}^\top \mathbf{X})^{-1} \mathbf{X}^T \mathbf{X} = \mathbf{X}$

3.5.2 Symmetric: $\mathbf{H}^T = \mathbf{H}$

Symmetric, since $\mathbf{H}^T = (\mathbf{X} (\mathbf{X}^\top \mathbf{X})^{-1} \mathbf{X}^T)^T = \mathbf{X} (\mathbf{X}^\top \mathbf{X})^{-1} \mathbf{X}^T = \mathbf{H}$

3.5.3 Idempotent: $\mathbf{H}\mathbf{H} = \mathbf{H}\mathbf{H}^T = \mathbf{H}^\top \mathbf{H} = \mathbf{H}$

Idempotent, i.e. $\mathbf{H}\mathbf{H} = \mathbf{H}\mathbf{H}^T = \mathbf{H}^\top \mathbf{H} = \mathbf{H}$. Indeed

$$\mathbf{H}\mathbf{H} = \mathbf{X} (\mathbf{X}^\top \mathbf{X})^{-1} \mathbf{X}^T \mathbf{X} (\mathbf{X}^\top \mathbf{X})^{-1} \mathbf{X}^T = \mathbf{X} (\mathbf{X}^\top \mathbf{X})^{-1} \mathbf{X}^T = \mathbf{H}$$

3.5.4 $\mathbf{H}(\mathbf{I} - \mathbf{H}) = \mathbf{0}$

This also implies that $\mathbf{H}(\mathbf{I} - \mathbf{H}) = \mathbf{0}_{n \times n}$

3.5.5 $(\mathbf{I} - \mathbf{H})$ is also symmetric and idempotent

$$(\mathbf{I} - \mathbf{H})(\mathbf{I} - \mathbf{H}) = (\mathbf{I} - \mathbf{H})^T(\mathbf{I} - \mathbf{H}) = (\mathbf{I} - \mathbf{H})(\mathbf{I} - \mathbf{H})^T = \mathbf{I} - \mathbf{H}$$

3.5.6 $\text{trace}(\mathbf{H}) = p$

$\text{trace}(\mathbf{H}) = p$, the number of LS coefficients we estimated.

3.6 Geometric Representation of LS

3.6.1 Estimation Space

- The columns of \mathbf{X} span a p -dimensional subspace in \mathbb{R}^n . This is a subspace that consists of vectors that can be written as linear combinations of the columns of X .

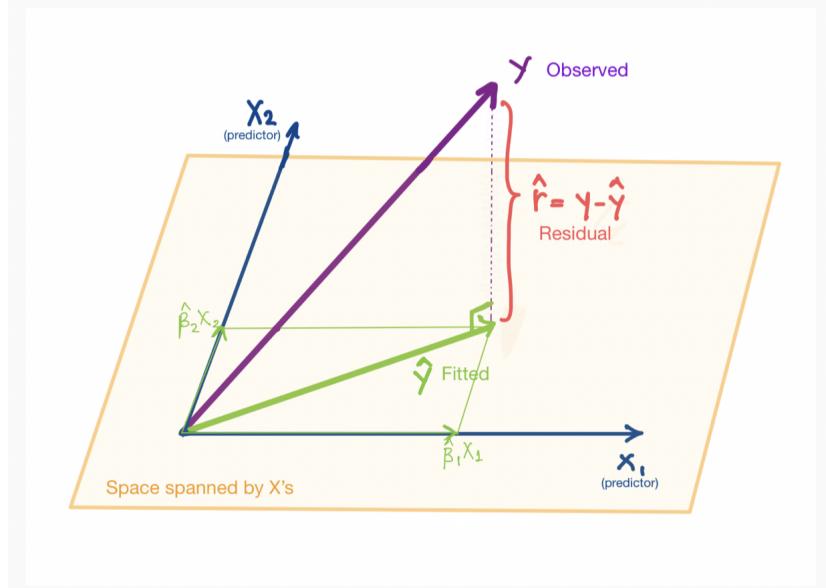


Figure 3.1

- The LS squares estimator $\hat{\beta}$ is obtained by minimizing the Euclidean distance between the vectors \mathbf{y} and $\hat{\mathbf{y}}$, i.e. $\|\mathbf{y} - \hat{\mathbf{y}}\|^2$. $\hat{\mathbf{y}}$ is the projection of \mathbf{y} onto the estimation space.
- $\mathbf{H}_{n \times n}$, projection/hat matrix is symmetric, unique, and idempotent.

3.6.2 Error Space

- The error space is an $(n - p)$ -dimensional space that is orthogonal to the estimation space. The projection matrix of the error space is $(\mathbf{I} - \mathbf{H})$.
- The residual \mathbf{r} is the projection of \mathbf{y} onto the error space, orthogonal to the estimation space. So, \mathbf{r} is orthogonal to any vector in the estimation space, including each column of X .
- When the intercept is included in the model, then

$$\sum_{i=1}^n r_i = 0$$

In general, $\sum_{i=1}^n r_i X_{ij} = 0, j = 1, \dots, p$ due to the normal equations.

3.7 Coefficient of determination, R-Square

$$R^2 = 1 - \frac{\sum_i (\hat{y}_i - y_i)^2}{\sum_i (y_i - \bar{y})^2} = 1 - \frac{RSS}{TSS}$$

An equivalent definition is

$$R^2 = \frac{\sum_i (\hat{y}_i - \bar{y})^2}{\sum_i (y_i - \bar{y})^2}$$

$$0 \leq R^2 \leq 1$$

3.8 Properties of LS Estimators

3.8.1 Unbiased: $\mathbb{E}(\hat{\beta}) = \beta$

$$\begin{aligned}\mathbb{E}(\hat{\beta}) &= \mathbb{E}((\mathbf{X}^T \mathbf{X})^{-1} \mathbf{X}^T y) = (\mathbf{X}^T \mathbf{X})^{-1} \mathbf{X}^T \mathbb{E}(y) \\ &= (\mathbf{X}^T \mathbf{X})^{-1} \mathbf{X}^T \mathbf{X} \beta = \beta\end{aligned}$$

3.8.2 Variance-Covariance Matrix of $\hat{\beta}$: $\text{Cov}(\hat{\beta}) = \sigma^2 (\mathbf{X}^T \mathbf{X})^{-1}$

$$\begin{aligned}\text{Cov}(\hat{\beta}) &= (\mathbf{X}^T \mathbf{X})^{-1} \mathbf{X}^T \text{Cov}(\mathbf{y}) \left((\mathbf{X}^\top \mathbf{X})^{-1} \mathbf{X}^\top \right)^\top \\ &= (\mathbf{X}^T \mathbf{X})^{-1} \mathbf{X}^T \sigma^2 \mathbf{X} \left(\mathbf{X}^\top \mathbf{X} \right)^{-1} \\ &= \sigma^2 (\mathbf{X}^T \mathbf{X})^{-1} \mathbf{X}^\top \mathbf{X} (\mathbf{X}^T \mathbf{X})^{-1} \\ &= \sigma^2 (\mathbf{X}^T \mathbf{X})^{-1}\end{aligned}$$

$$se(\hat{\beta}_i) = \hat{\sigma} \sqrt{((\mathbf{X}^T \mathbf{X})^{-1})_{ii}}$$

3.8.3 Properties of \hat{y} : $\mathbb{E}(\hat{y}) = \mathbf{X}\beta$, $Cov(\hat{y}) = \sigma^2 \mathbf{H}$

$$\mathbb{E}(\hat{y}) = \mathbb{E}(\mathbf{X} (\mathbf{X}^T \mathbf{X})^{-1} \mathbf{X}^\top y) = \mathbf{X} (\mathbf{X}^T \mathbf{X})^{-1} \mathbf{X}^\top \mathbf{X} \beta = \mathbf{X} \beta$$

$$Cov(\hat{y}) = Cov(\mathbf{H}y) = \mathbf{H}Cov(y)\mathbf{H}^T = \sigma^2 \mathbf{H} \mathbf{H}^T = \sigma^2 \mathbf{H}$$

3.8.4 Properties of \mathbf{r} : $\mathbb{E}(\mathbf{r}) = 0$, $Cov(\mathbf{r}) = \sigma^2 (\mathbf{I}_n - \mathbf{H})$

$$\mathbb{E}(\mathbf{r}) = \mathbb{E}(y - \hat{y}) = 0$$

$$Cov(\mathbf{r}) = Cov(y - \hat{y}) = Cov((\mathbf{I} - \mathbf{H})y) = \sigma^2 (\mathbf{I} - \mathbf{H})(\mathbf{I} - \mathbf{H})^T = \sigma^2 (\mathbf{I} - \mathbf{H})$$

3.8.5 Unbiased Variance Estimation $\mathbb{E}(\hat{\sigma}^2) = \sigma^2$

$$\mathbb{E}(\hat{\sigma}^2) = \frac{1}{n-p} \mathbb{E}(\mathbf{r}^T \mathbf{r}) = \frac{1}{n-p} \sigma^2(n-p) = \sigma^2$$

$n-p$: there are p elements in the diagonal of H are 1 and others are 0, so the diagonal of $\mathbf{I}_n - \mathbf{H}$ has $n-p$ terms with value 1. Then there are only $n-p$ terms with value $r_i^2 = \sigma^2$.

$$\frac{\mathbf{r}^T \mathbf{r}}{\sigma^2} = \frac{RSS}{\sigma^2} \sim \chi_{n-p}^2$$

3.9 The Gauss-Markov Theorem: LS estimator is the **BLUE**(best linear unbiased estimator)

If the errors are 1. Mean zero, 2. uncorrelated, 3. homoscedastic, the LS estimators have **the lowest variance within the class of linear estimators.**

Suppose we are interested in estimating a linear combination of β of the form:

$$\theta = \mathbf{c}^T \beta = \sum_{j=1}^p c_j \beta_j$$

LS estimators:

$$\hat{\theta}_{LS} = \mathbf{c}^T \hat{\beta} = \mathbf{c}^T (\mathbf{X}^T \mathbf{X})^{-1} \mathbf{X}^T y$$

This is a **linear** (linear combination of y_1, y_2, \dots, y_n) and **unbiased estimator** of θ . Its mean square error can be calculated as:

$$MSE(\hat{\theta}_{LS}) = \mathbb{E}(\hat{\theta}_{LS} - \theta)^2 = Var(\hat{\theta}_{LS})$$

Theorem 3.1 (Gauss-Markov Theorem)

$\hat{\theta}_{LS} = \mathbf{c}^T \hat{\beta}$ is the **BLUE**(best linear unbiased estimator) of the parameter $\mathbf{c}^T \beta$ for any vector $\mathbf{c} \in \mathbb{R}^p$



3.10 Maximum Likelihood Estimation, Distribution of LS estimates

3.10.1 Assumption $\mathbf{y} \sim \mathbf{N}_n(\mathbf{X}\beta, \sigma^2 \mathbf{I})$

Recall the normality assumption for the regression model:

$$y_i = \mathbf{x}_i^T \beta + \varepsilon_i, \quad i = 1, \dots, n, \text{ with } \varepsilon_i \sim N(0, \sigma^2)$$

This implies that $\mathbf{y} \sim \mathbf{N}_n(\mathbf{X}\beta, \sigma^2 \mathbf{I})$

3.10.2 LS estimator is the Maximum Likelihood Estimator (MLE)

We can show that the likelihood function can be written as:

$$L(\beta, \sigma^2 | \mathbf{y}) \propto \frac{RSS^{-\frac{n}{2}}}{n}$$

The value of β that maximizes the Likelihood function is *the Maximum Likelihood Estimator (MLE) of β* . This estimator is equal to the LS estimate of β .

3.10.3 Estimations $\hat{\beta} \sim N_p \left(\beta, \sigma^2 (\mathbf{X}^\top \mathbf{X})^{-1} \right)$, $\hat{\mathbf{y}} \sim N_n (\mathbf{X}\beta, \sigma^2 \mathbf{H})$, $\hat{\mathbf{r}} \sim N_n (\mathbf{0}, \sigma^2 (\mathbf{I}_n - \mathbf{H}))$

Recall the assumption for the linear regression model:

$$\mathbf{y} \sim N_n (\mathbf{X}\beta, \sigma^2 \mathbf{I})$$

Any affine transformation of \mathbf{y} will also have a Normal distribution ². We can show that:

$$\begin{aligned} \hat{\beta} &= (\mathbf{X}^\top \mathbf{X})^{-1} \mathbf{X}^\top \mathbf{y} \sim N_p \left(\beta, \sigma^2 (\mathbf{X}^\top \mathbf{X})^{-1} \right) \\ \hat{\mathbf{y}} &= \mathbf{H}\mathbf{y} \sim N_n (\mathbf{X}\beta, \sigma^2 \mathbf{H}) \\ \hat{\mathbf{r}} &= (\mathbf{I}_n - \mathbf{H})\mathbf{y} \sim N_n (\mathbf{0}, \sigma^2 (\mathbf{I}_n - \mathbf{H})) \end{aligned}$$

3.11 r Residuals' Properties

3.11.1 $n - p$ dimensions: $\mathbf{r} \in \mathbb{R}^{n-p}$

Although \mathbf{r} is a vector of dimension n , it always lies in a subspace of dimension $(n - p)$.

3.11.2 Distribution: $\hat{\sigma}^2 \sim \sigma^2 \frac{\chi_{n-p}^2}{n-p}$

\mathbf{r} behaves like a random vector with a distribution $N_{n-p} (\mathbf{0}, \sigma^2 \mathbf{I}_{n-p})$, so we have:

$$\hat{\sigma}^2 = \frac{\|\mathbf{r}\|^2}{n-p} \sim \sigma^2 \frac{\chi_{n-p}^2}{n-p}$$

3.11.3 $\hat{\mathbf{y}}$ and \mathbf{r} are independent

It can be show that $\hat{\mathbf{y}}$ and \mathbf{r} are uncorrelated since they are in orthogonal spaces. Since they also have a joint normal distribution, they are independent.

3.12 Testing Predictors (Coefficients)

3.12.1 Testing a Single Predictor $H_0 : \beta_j = 0$: t -test

Suppose you have a p predictors in your regression model and you want to test the hypothesis:

$$H_0 : \beta_j = c \text{ vs. } H_\alpha : \beta_j \neq c$$

- The t-test statistic we use is:

$$t = \frac{\hat{\beta}_j - c}{se(\hat{\beta}_j)} = \frac{\hat{\beta}_j - c}{\hat{\sigma} \sqrt{[(\mathbf{X}^\top \mathbf{X})^{-1}]_{jj}}} \sim T_{n-p}$$

under the null hypothesis H_0 .

- p-value = $2 \times$ the area under the curve of a T_{n-p} distribution more extreme than the observed statistic.

- The p -value returned by the *Im* function command is for $c = 0$.

3.12.2 Review the degree of freedom

The degrees of freedom of a t -test are determined by the denominator of the estimated variance $\hat{\sigma}^2$. Consider the following situations:

- In STAT 400: Test for $\theta = \alpha$, where $Z_1, \dots, Z_n \sim \mathcal{N}(\theta, \sigma^2)$

$$\frac{\hat{\theta} - \alpha}{se(\hat{\theta})} = \frac{\bar{Z} - \alpha}{\sqrt{\hat{\sigma}^2/n}} \sim T_{n-1}, \quad \hat{\sigma}^2 = \frac{\sum_i (Z_i - \bar{Z})^2}{n-1}$$

- In SLR: Test for $\beta_1 = c$, we have

$$\frac{\hat{\beta}_1 - c}{se(\hat{\beta}_1)} = \frac{\hat{\beta}_1 - c}{\hat{\sigma}/\sqrt{S_{XX}}} \sim T_{n-2}, \quad \hat{\sigma}^2 = \frac{RSS}{n-2}$$

- In MLR with p predictors (including the intercept): Test for $\beta_j = c$,

$$\frac{\hat{\beta}_j - c}{se(\hat{\beta}_j)} = \frac{\hat{\beta}_j - c}{\hat{\sigma} \sqrt{[(\mathbf{X}^\top \mathbf{X})^{-1}]_{jj}}} \sim T_{n-p}, \quad \hat{\sigma}^2 = \frac{RSS}{n-p}$$

3.12.3 Testing all predictors: F -test

Testing all predictors

$$\begin{cases} H_0 : \beta_2 = \beta_3 = \dots = \beta_p = 0 \\ H_a : \beta_j \neq 0, \quad \text{for some } j, j = 2, \dots, p \end{cases}$$

- Under the Null hypothesis, the test statistic:

$$\begin{aligned} F &= \frac{FSS(X_2, \dots, X_p)}{p-1} \div \frac{RSS(X_2, \dots, X_p)}{n-p} \\ &= \frac{MS(Reg)}{MS(\text{Error})} \sim F_{p-1, n-p} \end{aligned}$$

Large values of F lead to conclusion H_α .

- This is the overall F test of whether or not there is a regression relation between the response variable Y and the set of X variables.

Source	df	SS	MS	F-test
Regression	$p - 1$	FSS	$FSS/(p - 1)$	$MS(\text{Reg})/\text{MSE}$
Error	$n - p$	RSS	$RSS/(n - p)$	
Total	$n - 1$	TSS		

Figure 3.2

3.12.4 Partial F -test

In general, consider the following partition of the design matrix into two sub-matrices \mathbf{X}_1 and \mathbf{X}_2 , that is

$$\mathbf{X}_{n \times p} = (\mathbf{X}_{1n \times (p-q)}, \mathbf{X}_{2n \times q})$$

The corresponding partition of the regression parameter is:

$$\beta^T = (\beta_1^T, \beta_2^T)$$

where β_1 is $(p - q) \times 1$ and β_2 is $q \times 1$

This partition is used to test the hypothesis:

$$\begin{cases} H_0 : \beta_2 = \mathbf{0}, \text{ i.e., } \mathbf{y} = \mathbf{X}_1\beta_1 + \text{error} \\ H_\alpha : \beta_2 \neq \mathbf{0}, \text{ i.e., } \mathbf{y} = \mathbf{X}_1\beta_1 + \mathbf{X}_2\beta_2 + \text{error} \end{cases}$$

To test this hypothesis, the test statistic is:

$$F = \frac{(RSS_0 - RSS_\alpha)/q}{RSS_\alpha/(n-p)} \sim F_{q, n-p}$$

where $RSS_0 = \text{Residual sum of squares for the model under } H_0$; $RSS_\alpha = \text{Residual sum of squares for the model under } H_\alpha$

Numerator: variation in the data not explained by the reduced model, but explained by the full model.

Denominator: variation in the data not explained by the full model (i.e., not explained by either model), which is used to estimate the error variance.

Reject H_0 , if F test statistic is large, that is, the variation missed by the reduced model, when being compared with the error variance, is significantly large.

F-test statistic can be rewritten:

$$F = \frac{(RSS_0 - RSS_\alpha)/q}{RSS_\alpha/(n-p)} = \frac{((1-R^2_R) - (1-R^2_F))/(df_F - df_R)}{(1-R^2_F)/df_F} = \frac{(R^2_F - R^2_R)/(df_R - df_F)}{(1-R^2_F)/df_F}$$

Note that this test statistic is not appropriate when the full and reduced regression models do not contain the intercept term β_0 .

3.13 Permutation Tests (When the normal distribution hypothesis doesn't hold)

The distribution of the data is *unknown*. - A test statistic is a function of the data; denote it $g(\text{ data })$.

- The test statistic tends to take extreme values under the alternative hypothesis H_α .

3.13.1 Procedure

Procedure to conduct a permutation test

1. Form the test statistic $g(\text{ data })$ which tends to take extreme values under the alternative hypothesis.
2. Evaluate the test statistic on the observed data, denoted by g_0 .
3. Find the distribution of $g(\text{ data })$, when data are generated from H_0 .
4. Calculate the p -value, that is the following probability:

$$\mathbb{P}(g(\text{ data }) \text{ is more extreme than the observed } g_0 \mid \text{ data } \sim H_0)$$

3.13.2 Calculation of the p-value: Monte Carlo method

We can obtain an approximation of $\mathbb{E}(Y)$ as follows:

1. Generate $N = 1000$ samples from this distribution, Y_1, \dots, Y_N
2. Approximate the mean by

$$\mathbb{E}(Y) \approx \frac{1}{N} \sum_{i=1}^N Y_i$$

That is, population mean \approx sample mean (when N is large).

This method also works if we want to approximate the expected value of a *function* of a random variable:

$$\mathbb{E}(f(Y)) \approx \frac{1}{N} \sum_{i=1}^N f(Y_i)$$

3.14 Confidence Intervals for β_j , Confidence Region for β

3.14.1 Confidence Intervals for β_j

Recall that

$$\hat{\beta} = (\mathbf{X}^\top \mathbf{X})^{-1} \mathbf{X}^\top \mathbf{y} \sim \mathbf{N}_{\mathbf{P}} \left(\beta, \sigma^2 (\mathbf{X}^\top \mathbf{X})^{-1} \right)$$

An $(1 - \alpha)100\%$ CI(Confidence Interval) for β_j can be written as

$$(\hat{\beta}_j \pm T_{n-p}(\frac{\alpha}{2}) se(\hat{\beta}_j)) = (\hat{\beta}_j \pm T_{n-p}(\frac{\alpha}{2}) \hat{\sigma} \sqrt{[(\mathbf{X}^\top \mathbf{X})^{-1}]_{jj}})$$

Justification: The vector β 's confidence interval is a family/joint interval for all the betas, so it will be wider than the individual β_j 's intervals.

3.14.2 Confidence Region for β

β is the entire vector,

$$\beta - \hat{\beta} \sim \mathbf{N}_{\mathbf{P}} \left(0, \sigma^2 (\mathbf{X}^\top \mathbf{X})^{-1} \right)$$

Thus the quadratic form:

$$\frac{(\beta - \hat{\beta})^\top \mathbf{X}^\top \mathbf{X} (\beta - \hat{\beta})}{p \hat{\sigma}^2} \sim F_{p, n-p}$$

We can construct a $(1 - \alpha)100\%$ confidence region for β to be all the points in the following ellipsoid,

$$\frac{(\beta - \hat{\beta})^\top \mathbf{X}^\top \mathbf{X} (\beta - \hat{\beta})}{p \hat{\sigma}^2} < F(\alpha; p, n-p)$$

Where $F(\alpha; p, n-p)$ is defined to be the point such that

$$\mathbb{P}(F_{p, n-p} > F(\alpha; p, n-p)) = \alpha$$

3.15 Confidence/Prediction Intervals for New Observations

Set $\mathbb{E}(Y|x^*) = \mu^* = (x^*)^\top \beta$

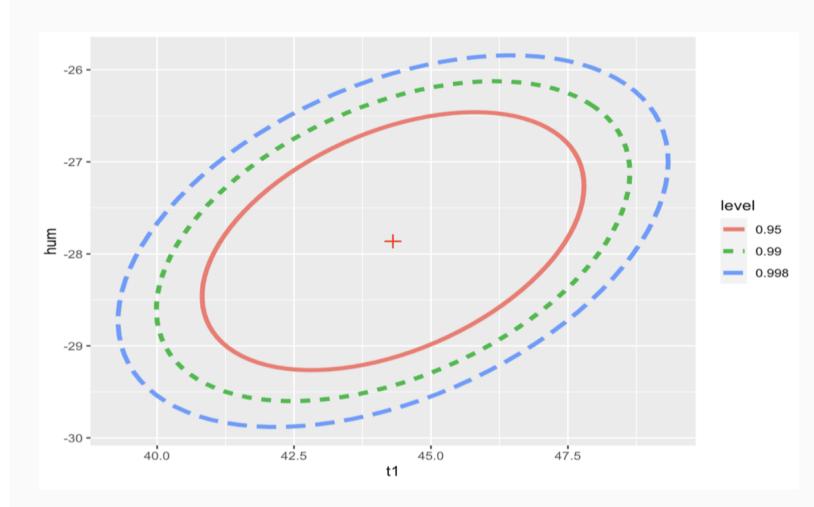


Figure 3.3

3.15.1 Confidence Interval for $\mu^* = (x^*)^T \beta$

The Gauss-Markov theorem tells us that the BLUE (Best Linear Unbiased Estimate) of μ^* is:

$$\hat{\mu}^* = (x^*)^T \hat{\beta} = (x^*)^T (X^T X)^{-1} X^T y$$

Then,

$$\begin{aligned} \mathbb{E}[(\hat{\mu}^* - \mu^*)^2] &= Var(\hat{\mu}^*) \\ &= \sigma^2 (x^*)^T (X^T X)^{-1} x^* \\ se(\hat{\mu}^*) &= \hat{\sigma} \sqrt{(x^*)^T (X^T X)^{-1} x^*} \end{aligned}$$

Where

$$\frac{\hat{\mu}^* - \mu}{se(\hat{\mu}^*)} \sim t(n-p)$$

A $(1 - \alpha)100\%$ CI (confidence interval) for μ^* is:

$$\hat{\mu}^* \pm T_{n-p}(\frac{\alpha}{2}) se(\hat{\mu}^*)$$

3.15.2 Prediction Interval for $y^* = (x^*)^T \beta + e^*$

The best estimate for y^* at a future observation x^* is also

$$\hat{y}^* = (x^*)^T \hat{\beta}$$

Then,

$$\begin{aligned}
 Var(\hat{y}^*) &= \mathbb{E}[((x^*)^T \hat{\beta} - y^*)^2] \\
 &= \mathbb{E}[((x^*)^T \hat{\beta} - ((x^*)^T \beta + e^*))^2] \\
 &= \mathbb{E}[((x^*)^T \hat{\beta} - ((x^*)^T \beta)^2] + \mathbb{E}[(e^*)^2] \\
 &= Var(\hat{\mu}^*) + Var(e) \\
 &= \sigma^2[1 + (x^*)^T (X^T X)^{-1} x^*] \\
 se(\hat{y}^*) &= \hat{\sigma} \sqrt{1 + (x^*)^T (X^T X)^{-1} x^*}
 \end{aligned}$$

Where

$$\frac{\hat{y}^* - y^*}{se(\hat{y}^*)} \sim t(n - p)$$

A $(1 - \alpha)100\%$ PI (prediction interval) for y^* is:

$$\hat{y}^* \pm T_{n-p}(\frac{\alpha}{2}) se(\hat{y}^*)$$

3.15.3 Mahalanobis distance

$$\mathbf{X}_{n \times p} = \begin{bmatrix} \mathbf{x}_1 & \mathbf{x}_2 & \cdots & \mathbf{x}_n \end{bmatrix}^T$$

For any observation vector $\mathbf{x}_{p \times 1} = \begin{pmatrix} 1 \\ \mathbf{z} \end{pmatrix}$, where \mathbf{z} denotes the value of predictors without the intercept.

We would like to use **Mahalanobis distance** to quantify the distance between observation vector $\mathbf{x}_{p \times 1}$ and its sample meaning $\bar{\mathbf{x}}$.

The sample covariance matrix of $\mathbf{Z} = [\mathbf{z}_1 \ \mathbf{z}_2 \ \cdots \ \mathbf{z}_n]$ is:

$$\hat{\Sigma}_{(p-1) \times (p-1)} = \frac{1}{n-1} \sum_{i=1}^n (\mathbf{z}_i - \bar{\mathbf{z}})(\mathbf{z}_i - \bar{\mathbf{z}})^T$$

Then the $(x^*)^T (X^T X)^{-1} x^*$ can be written as

$$(x^*)^T (X^T X)^{-1} x^* = \frac{1}{n} + \frac{1}{n-1} (\mathbf{z}^* - \bar{\mathbf{z}})^T \hat{\Sigma}^{-1} (\mathbf{z}^* - \bar{\mathbf{z}})$$

The second term in the right hand side ($\frac{1}{n-1} (\mathbf{z}^* - \bar{\mathbf{z}})^T \hat{\Sigma}^{-1} (\mathbf{z}^* - \bar{\mathbf{z}})$) is the so-called Mahalanobis distance from \mathbf{z}^* to the center of the data $\bar{\mathbf{z}}$ (the sample mean).

Then we can write,

$$\begin{aligned} \text{se}(\hat{\mu}^*) &= \hat{\sigma} \sqrt{\mathbf{x}^{*\top} (\mathbf{X}^\top \mathbf{X})^{-1} \mathbf{x}^*} \\ &= \hat{\sigma} \sqrt{\frac{1}{n} + \frac{1}{n-1} (\mathbf{z}^* - \bar{\mathbf{z}})^T \hat{\Sigma}^{-1} (\mathbf{z}^* - \bar{\mathbf{z}})} \\ \text{se}(\hat{y}^*) &= \hat{\sigma} \sqrt{1 + (\mathbf{x}^*)^\top (\mathbf{X}^\top \mathbf{X})^{-1} \mathbf{x}^*} \\ &= \hat{\sigma} \sqrt{1 + \frac{1}{n} + \frac{1}{n-1} (\mathbf{z}^* - \bar{\mathbf{z}})^T \hat{\Sigma}^{-1} (\mathbf{z}^* - \bar{\mathbf{z}})} \end{aligned}$$

Since $\text{se}(\hat{y}^*)$ has an extra 1, when the sample size n goes to infinity,

$$\text{se}(\hat{\mu}^*) \rightarrow 0$$

$$\text{se}(\hat{y}^*) \rightarrow \sigma$$

Chapter 4 MLR: unusual observations

Recall, that we can write the MLR model as:

$$\mathbf{y} \sim \mathcal{N}(\mathbf{X}\beta, \sigma^2 \mathbf{I}_n)$$

- Error: assumed to be iid, $\varepsilon_i \sim \mathcal{N}(0, \sigma^2)$
- Model: assumed to be linear in the parameters, i.e., $\mathbb{E}(\mathbf{y}) = \mathbf{X}\beta$

We might have unusual observations.

4.1 1. High leverage points: $h_i \geq \frac{2p}{n}$

4.1.1 Leverage Points

The diagonal elements of $\mathbf{H} = \mathbf{X}(\mathbf{X}^T \mathbf{X})^{-1} \mathbf{X}^\top$,

$$h_i = H_{ii} = \mathbf{x}_i (\mathbf{X}^T \mathbf{X})^{-1} \mathbf{x}_i^\top = \frac{\text{Var}(x_i^T \hat{\beta})}{\sigma^2}$$

are called **leverages** and are very useful diagnostics. h_i gives a measure (invariant under any affine transformation of \mathbf{X}) of how far the i -th observation is from the center of the data (in the X -space).

For simple linear regression:

$$h_i = \frac{1}{n} + \frac{(x_i - \bar{x})^2}{\sum_i (x_i - \bar{x})^2}$$

In general:

$$\begin{aligned} h_i &= \mathbf{x}_i^\top (\mathbf{X}^\top \mathbf{X})^{-1} \mathbf{x}_i \\ &= \frac{1}{n} + \frac{1}{n-1} (\mathbf{z}_i - \bar{\mathbf{z}})^T \hat{\Sigma}^{-1} (\mathbf{z}_i - \bar{\mathbf{z}}) \end{aligned}$$

where

$$\hat{\Sigma}_{(p-1) \times (p-1)}^{-1} = \frac{1}{n-1} \sum_{i=1}^n (\mathbf{z}_i - \bar{\mathbf{z}})(\mathbf{z}_i - \bar{\mathbf{z}})^T$$

is the sample covariance of the $(p-1)$ predictor variables. The second term in the right hand side $(\frac{1}{n-1} (\mathbf{z}_i - \bar{\mathbf{z}})^T \hat{\Sigma}^{-1} (\mathbf{z}_i - \bar{\mathbf{z}}))$ is the so-called **Mahalanobis distance** from \mathbf{z}_i to the data center $\bar{\mathbf{z}}$

4.1.2 Properties of the Leverage: $0 < h_i < 1, \sum_i h_i = p$

Recall that the hat matrix is idempotent $\mathbf{H} = \mathbf{HH}^\top$ and has $\text{tr}(\mathbf{H}) = p$

These imply that

$$\sum_i h_i = p \text{ and } \sum_j H_{ij}^2 = h_i$$

For a given i we can decompose the last sum as follows:

$$\begin{aligned} \sum_j H_{ij}^2 &= H_{ii}^2 + \sum_{i \neq j} H_{ij}^2 = h_i \\ \Rightarrow \sum_{i \neq j} H_{ij}^2 &= h_i(1 - h_i) \Rightarrow h_i(1 - h_i) > 0 \end{aligned}$$

From this we can conclude the following properties of h_i :

$$0 < h_i < 1, \quad \sum_i h_i = p$$

4.1.3 Fitted Values and Leverage: $\text{Var}(\hat{y}_i) = \sigma^2 h_i, \text{Var}(r_i) = \sigma^2(1 - h_i)$

Recall the equation $\hat{\mathbf{y}} = \mathbf{Hy}$.

$$\begin{aligned} \hat{y}_i &= \sum_{t=1}^n H_{it} y_t \\ &= h_i y_i + \sum_{t \neq i} H_{it} y_t \end{aligned}$$

This means that $h_i = \frac{d\hat{y}_i}{dy_i}$

When h_i is **large (close to 1)**, \hat{y}_i relies heavily on y_i (instead of using the information from other data points), therefore \hat{y}_i will be “forced” to be **close** to the observed y_i .

Consequently, the variance for the residual r_i will be small, and the variance for the fit \hat{y}_i will be large (since the fit from another data set would be quite different).

$$\text{Var}(\hat{y}_i) = \sigma^2 h_i, \text{Var}(r_i) = \sigma^2(1 - h_i)$$

4.1.4 High-leverage Points: $h_i \geq \frac{2p}{n}$

Good high-leverage points: its y point follows the pattern of the rest of the data, but with an x_i value that is far away from the sample mean.

Bad high-leverage points: its y value does not follow the pattern suggested by the rest of the data; the LS fitting might change a lot if we remove this point.

4.2 Residuals: Standardized Residuals vs. Studentized residuals

The residuals $r_i = y_i - \hat{y}_i$ do not have a constant variance. So they need to be standardized.

4.2.1 Difference between ε and r

ε : true residuals (our theoretical quantities)

r : estimated residuals - Both residuals are normally distributed, but:

$$\varepsilon \sim \mathcal{N}_n (\mathbf{0}, \sigma^2 \mathbf{I}_n), \quad r \sim \mathcal{N}_n (\mathbf{0}, \sigma^2 (\mathbf{I}_n - \mathbf{H}))$$

where \mathbf{H} is the projection/hat matrix.

- The errors ε_i 's have equal variance and are independent, while the residuals r_i 's have unequal variance and are correlated.

$-\mathbb{E}(\varepsilon) = \mathbb{E}(r) = \mathbf{0}$. But

$$\sum_i \varepsilon_i \neq 0, \quad \sum_i r_i = 0$$

(by default we assume an intercept is included in the model)

4.2.2 Standardized Residuals: $r_i^* = \frac{r_i}{\hat{\sigma}\sqrt{1-h_i}}$

Since $r_i \sim \mathcal{N}(0, \sigma^2(1-h_i))$, it is reasonable to consider a standardization of the residuals in this form:

$$r_i^* = \frac{r_i}{\hat{\sigma}\sqrt{1-h_i}}, \quad i = 1, \dots, n$$

- $\sum_i r_i^*$ is no longer zero.
- Since the r_i is not independent of $\hat{\sigma}$, each r_i^* is **not distributed as a T distribution**.
- As an approximation, we can view the r_i^* 's as iid $\mathcal{N}(0, 1)$ random variables, although they are not Normally distributed and they are slightly correlated.

4.2.3 Studentized Residuals: $t_i = r_i^* \left(\frac{n-p-1}{n-p-r_i^{*2}} \right)^{1/2}$

- The studentized residuals are based on the idea of leave-one-out (also known as jackknife residuals).

- Here is the leave-one-out idea:

1. Run a regression model on the $(n-1)$ samples with the i -th sample (x_i, y_i) removed.
2. Denote the leave-one-out estimates of the regression coefficient and error variance by $\hat{\beta}_{(i)}$ and $\hat{\sigma}_{(i)}$, where the notation (i) means "excluding the i -th observation."
3. Then, check the discrepancy between observations y_i and the fitted value $\hat{y}_{(i)} = \mathbf{x}^T \hat{\beta}_{(i)}$

- Define the Studentized Residuals as:

$$t_i = \frac{y_i - \hat{y}(i)}{\hat{\sigma}_{(i)} \left(1 + x_i^T \left(\mathbf{X}_{(i)}^\top \mathbf{X}_{(i)} \right)^{-1} x_i \right)^{1/2}} = \frac{y_i - \hat{y}(i)}{\hat{\sigma}_{(i)} \sqrt{1 - h_i}}$$

which follows a T_{n-p-1} distribution if $y_i \sim \mathcal{N}(\mathbf{x}_i^T \beta, \sigma^2)$

- One can also show that r_i^* and t_i are a monotone transformation of each other.

- We do not need to run the model n times to get the estimates $\hat{\beta}_{(i)}$ and $\hat{\sigma}_{(i)}$ since it can be shown that:

$$t_i = r_i^* \left(\frac{n-p-1}{n-p-r_i^{*2}} \right)^{1/2}$$

4.3 2. Outlier (Large $|r_i^*|$)

4.3.1 Outlier test

Outliers are observations that do not fit the model, but Outliers are not necessarily observations with large residuals.

We need to used the studentized residuals for the outlier test.

Under the Null hypothesis H_0 ,

$$t_i \sim T_{n-p-1}$$

We can use t-test to test the i^{th} observation: form a PI at x_i .

(an example of data snooping)

Large $|r_i^*| \Rightarrow$ Large $|t_i| \Rightarrow$ Reject Null hypothesis \Rightarrow Outlier

4.3.2 Bonferroni Correction

Suppose we are testing m hypothesis simultaneously.

For each test, we use a significant level α . That is, the chance to make a **overall** Type I error is α .

Suppose we want to control the overall type I error rate (for all m tests) to be 95%.

We should set the individual significance levels to be $\alpha = 5\% / m$

When we test outliers, since the T distribution is bipartite, we need the confidence level: $\alpha / (2 * n)$

4.3.3 What we should do with outliers?

Points should not be routinely deleted simply because they do not fit the model. No data snooping!

Outliers, as well as other unusual observations discussed here, often flag potential problems of the current

model. Instead of dropping them, maybe, try a new alternative model.

4.4 3. Highly Influential Points: Large $D_i = \frac{r_i^{*2}}{p} \left(\frac{h_i}{1-h_i} \right)$ ($D_i \geq 1$)

4.4.1 Influential observations

Observations whose removal greatly affects the regression analysis are called **influential observations**.

An **influential observations** may be (or may not) an **outlier** or a **high-leverage observation**; or may be both: an outlier and a high-leverage observation.

We will use the **Cook's distance** to detect influential observations.

$$D_i = \frac{\|\mathbf{X}\hat{\beta} - \mathbf{X}\hat{\beta}_{(i)}\|^2}{p\hat{\sigma}^2} = \frac{\|\hat{\mathbf{y}} - \hat{\mathbf{y}}_{(i)}\|^2}{p\hat{\sigma}^2} = \frac{r_i^{*2}}{p} \left(\frac{h_i}{1-h_i} \right)$$

which indicates that highly influential points are either outliers (large $|r_i^*|$) or high-leverage points (large h_i) or both.

A rule-of-thumb: observations with $D_i \geq 1$ are highly influential.

Chapter 5 MLR Diagnostics: Checking Assumptions

5.1 Classical Linear Model (CLM) assumption: Gauss-Markov Assumption +

$$\varepsilon \sim^{IID} \mathcal{N}(0, \sigma^2)$$

1. Constant Variance (Homoscedasticity)
2. Normality
3. Uncorrelated errors (No-Autocorrelation)
4. Linearity: $\mathbb{E}(y) = \mathbf{X}\beta$
5. Random Sampling
(1-5 calls Gauss-Markov Assumption)
6. $\mathbf{Y} = \beta\mathbf{X} + \varepsilon$, where $\varepsilon \sim^{IID} \mathcal{N}(0, \sigma^2)$
(1-6 calls Classical Linear Model (CLM) assumption)

5.2 Check 1. Constancy of Variance(Homoscedasticity)

5.2.1 Method 1: graph *residuals* against *Fitted Values* \hat{y}

SLR:

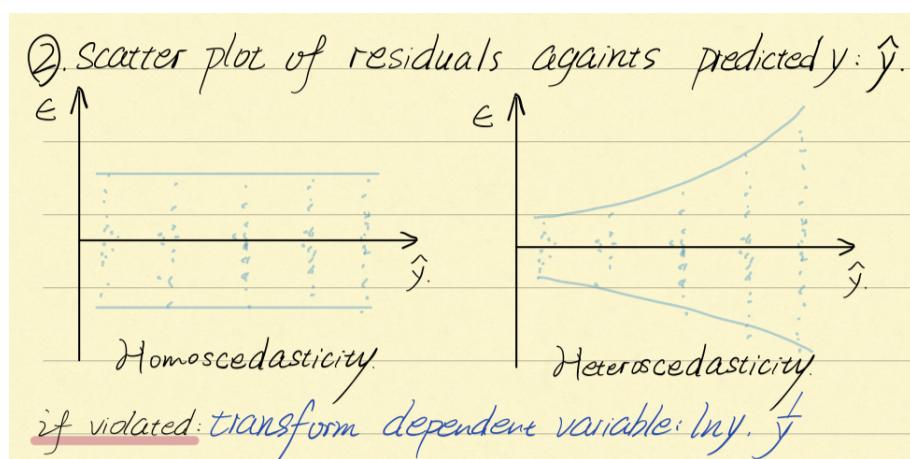


Figure 5.1

MLR:

If the variance is constant, the residuals will look like a football-shaped cloud. Check residual plots and look

for a “fan” type shape or trends.

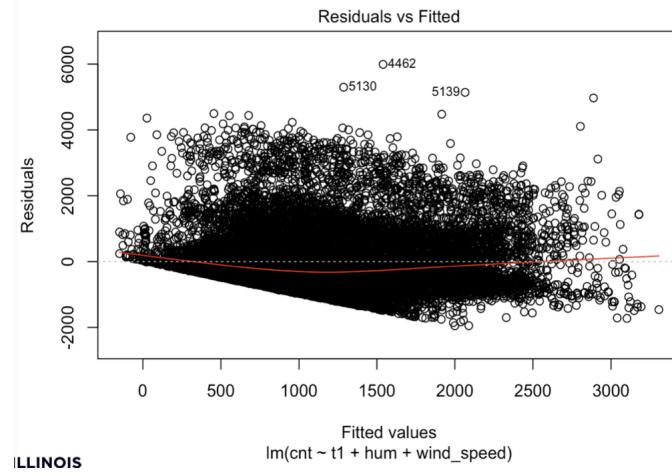


Figure 5.2

5.2.2 Method 2: Breusch-Pagan Test

$$\text{BP} = nR^2$$

where R^2 is the *coefficient of Determination* between the **squared residuals** r_i^2 of LS regression and the **covariates** (or a sub-set) X_1, X_2, \dots, X_p .

$$H_0 : \text{BP} \sim \chi_{p-1}^2 \text{ (asymptotically)}$$

5.2.3 What happen if Heteroscedasticity

Unchanges:

1. The estimator will still be *unbiased* ($\mathbb{E}(\hat{\beta}) = \beta$), *consistent* ($\lim_{n \rightarrow \infty} \hat{\beta} = \beta$), but *inefficient*.
2. Interpretation of R^2 is not changed.

Changes:

1. $\hat{\beta}$ will be *inefficient*.
2. invalidates variance formulas for OLS estimators ($\text{Var}(\hat{\beta}_i)$).
3. usual F -test, t -test are not valid.
4. *Gauss-Markov Theorem* doesn't hold: OLS is no longer the BLUE (best linear unbiased estimator).

5.2.4 What can we do if Heteroscedasticity

- (i) Take logarithms of each of the variables
- (ii) Use suitably modified standard errors
- (iii) Use a generalised least squares (GLS) procedure

5.2.5 Remedial measure: Variance Stabilizing Transformations \sqrt{Y} , $\log Y$, $\frac{1}{Y}$ or $\frac{1}{Y+1}$

SLR:

If violates Homoscedasticity: Transform *dependent variable*: $\ln y$, $\frac{1}{y}$

MLR:

Find a transformation of the response, $h(Y)$, to achieve constant variance.

How does it work?

- Suppose h is a smooth function.
- Using Taylor's theorem, the expansion of $h(Y)$ around $E(Y)$ is:

$$h(Y) = h(E(Y)) + h'(E(Y))(Y - E(Y)) + \text{Remainder}$$

- The remainder is assumed small with high probability and we can ignore it:

$$\text{Var}(h(Y)) \approx (h'(E(Y)))^2 \text{Var}(Y)$$

- We want to choose a transformation h such that $\text{Var}(h(Y))$ is approximately constant.

Example 1:

- Suppose that the variance of Y is proportional to the mean of Y , i.e., $\text{Var}(Y) \propto E(Y)$
- Select h such that:

$$h'(z) = \frac{1}{\sqrt{z}} \Rightarrow h(z) \propto \sqrt{z}$$

- When plugging-in the value of $h'(z)$ evaluated at $E(Y)$ in the variance of $h(Y)$ equation, the variance of $h(Y)$ will be approximately constant. Indeed,

$$\text{Var}(\sqrt{Y}) \approx \left(\frac{1}{\sqrt{E(Y)}} \right)^2 \text{Var}(Y) = \frac{\text{Var}(Y)}{E(Y)} \approx \text{const}$$

Example 2:

- Suppose that the variance of Y is proportional to the squared mean of Y , i.e., $\text{Var}(Y) \propto (E(Y))^2$.

- Select h such that:

$$h'(z) = \frac{1}{z} \Rightarrow h(z) = \log(z)$$

- Then,

$$\text{Var}(\log Y) \approx \frac{1}{(\text{E}(Y))^2} \text{Var}(Y) \approx \text{const}$$

Example 3:

$$\text{Var}(Y) \propto (\text{E}(Y))^4.$$

$$h(Y) = \frac{1}{Y} \text{ or } \frac{1}{Y+1}$$

5.3 Check 2. Normality

5.3.1 Method 1: Histogram, graph residuals against its frequency

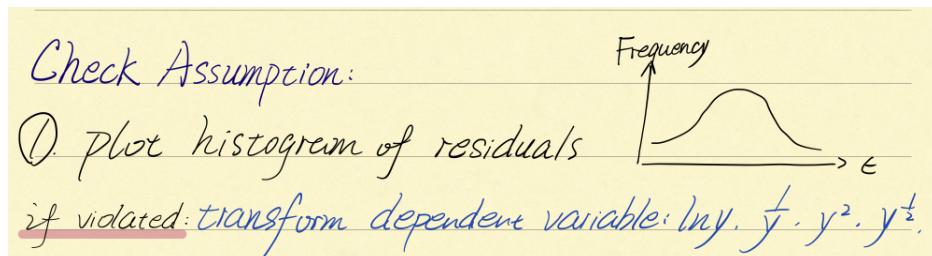


Figure 5.3

If violates Normality: Transform *dependent variable*: $\ln y, \frac{1}{y}, y^2, \sqrt{y}$

5.3.2 Method 2: QQ-Plot, graph residuals against its frequency

- Suppose that we have a sample z_1, z_2, \dots, z_n .

- We wish to examine the hypothesis that the z 's are a sample from a normal distribution with mean μ and variance σ^2 .

QQ-Plot:

1. Order the z 's: $z_{(1)}, z_{(2)}, \dots, z_{(n)}$.
2. Compute $u_i = \Phi^{-1} \left(\frac{i}{n+1} \right)$, where Φ is the cdf of the $N(0, 1)$ and i is the order if the data ($i = 1, 2, \dots, n$).

3. Plot $z_{(i)}$ against u_i .

⇒ If the z 's are normal, the plot should be approximately a straight line.

```
plot(bikeshare.ml, which=2)
```

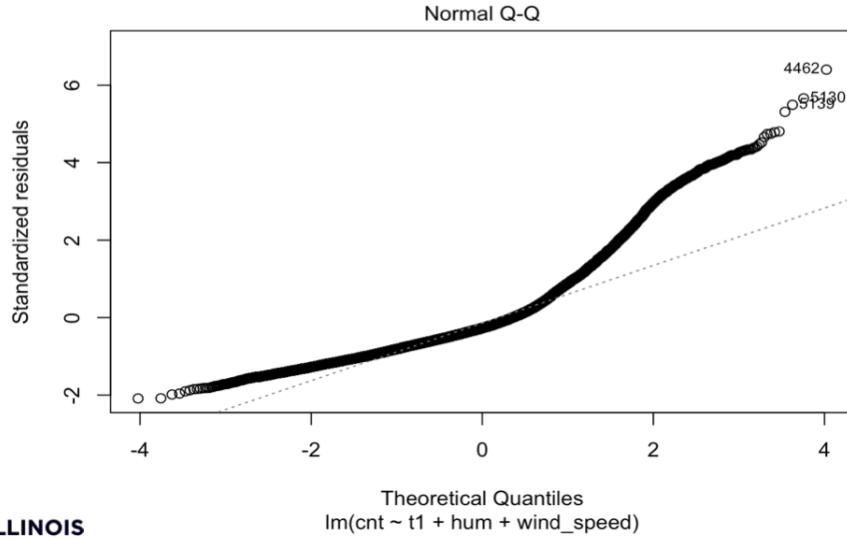


Figure 5.4

5.3.3 Method 3: Shapiro-Wilk Test (good for $n \leq 50$)

$$W = \frac{(\sum_{i=1}^n a_i r_{(i)})^2}{\sum_{i=1}^n (r_i - \bar{r})^2}$$

where $r_{(i)}$ is the i th largest value of the r_i 's and the a_i terms are calculated using the means, variances, and covariances of the r_i s.

Small values of W will lead to rejection of the null hypothesis.

5.3.4 Method 4: Kolmogorov-Smirnov Test (good for $n > 50$)

$$D_n = \sup_x |F_n(x) - \Phi(x)|$$

where $\Phi(x)$ is the cdf of the Normal and F_n the empirical distribution function F_n for n i.i.d. ordered observations X_i is defined as

$$F_n(x) = \frac{1}{n} \sum_{i=1}^n \mathbf{1}_{(-\infty, x]}(X_i)$$

Small values of D will lead to rejection of the null hypothesis.

5.3.5 Remedial measure: Box-Cox Transformations of Y

Suppose each $y_i > 0$, and consider the following transformation:

$$g_\lambda(y) = \begin{cases} \frac{y^\lambda - 1}{\lambda}, & \lambda \neq 0 \\ \log(y), & \lambda = 0 \end{cases}$$

Choose λ that maximizes the likelihood of the data, under the assumption that the transformed data $g_\lambda(\mathbf{y})$ has a normal distribution:

$$g_\lambda(\mathbf{y}) = \mathbf{X}\beta + \varepsilon, \quad \varepsilon \sim N_n(\mathbf{0}, \sigma^2 \mathbf{I})$$

- The log-likelihood function for $\lambda \neq 0$ is:

$$L(\lambda) = -\frac{n}{2} \log(RSS_\lambda/n) + (\lambda - 1) \sum_{i=1}^n \log(y_i)$$

where RSS_λ is the RSS when $g_\lambda(\mathbf{y})$ is the response, and for $\lambda = 0$ is:

$$L(0) = -\frac{n}{2} \log(RSS_0/n) - \sum_{i=1}^n \log(y_i)$$

The second term in these log-likelihood function corresponds to the Jacobian of the transformation.

In **R**, we can graph the log-likelihood as a function of $\lambda(L(\lambda))$ versus $\lambda \in (-2, 2)$ and then pick the maximizer $\hat{\lambda}$.

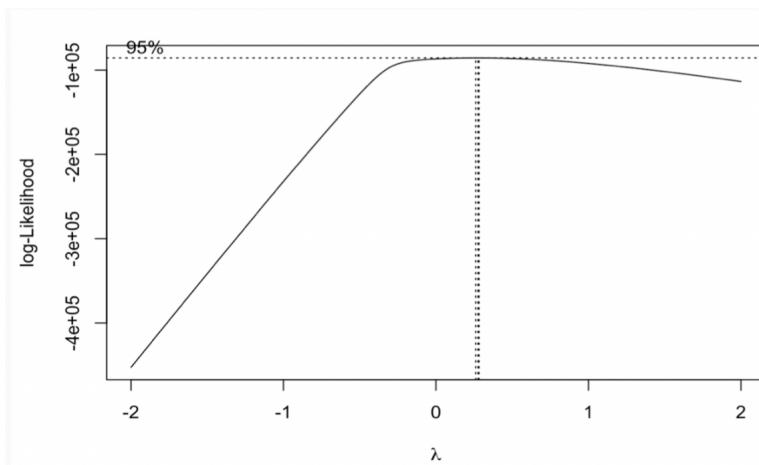


Figure 5.5

It is common to round $\hat{\lambda}$ to a nearby value like:

$-1, -0.5, 0, 0.5$, or 1

then the transformation defined by $\hat{\lambda}$ is easier to interpret.

To answer the question whether we really need the transformation g_λ , we can do hypothesis testing ($H_0 : \lambda = 1$), or equivalently construct a Confidence Interval for λ as follows:

$$\{\lambda : L(\lambda) > L(\hat{\lambda}) - \frac{1}{2}\chi_1^2(1 - \alpha)\}$$

5.4 Checking 3. Serial Dependence

5.4.1 Method 1: graph residuals against index variable(time or case number)

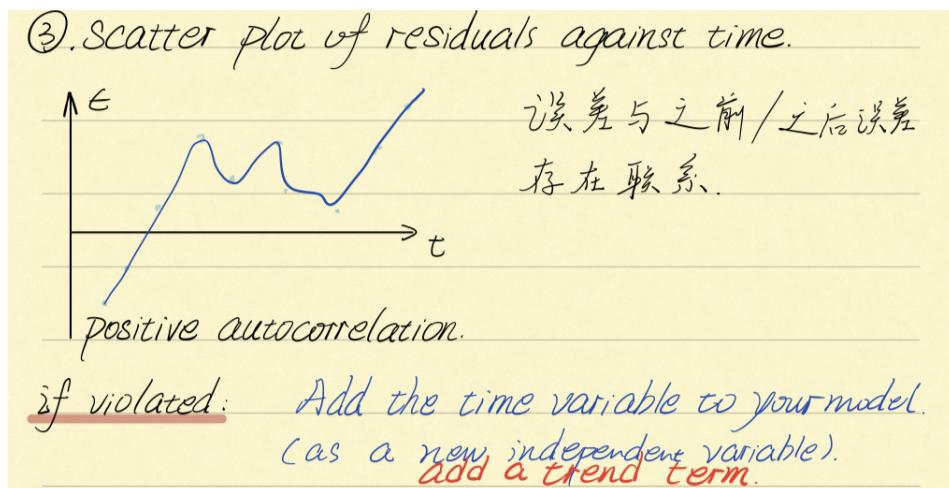


Figure 5.6

If violates No-Autocorrelation: add a new independent variable (t).

5.4.2 Method 2: Durbin Watson test

SLR:

$$d = \frac{\sum_{i=2}^n (e_i - e_{i-1})^2}{\sum_{i=1}^n e_i^2}$$

Compare d with d_L and d_U .

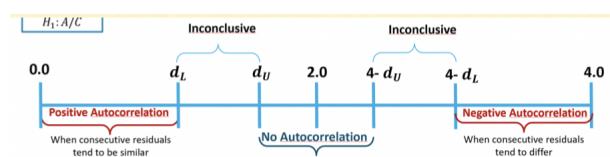


Figure 5.7

MLR:

$$DW = \frac{\sum_{k=1}^{n-1} (r_k - r_{k+1})^2}{\sum_{k=1}^n r_k^2}$$

if $DW < 2$, then there is evidence for positive serial dependence.

5.5 Checking 4. Non-Linearity

5.5.1 Method 1: Partial Regression Plots

We want to know **the relationship between the response Y and a predictor X_k** after the effect of the other predictors has been removed.

To remove the effect of the other predictors, run the following two regression models:

$$Y \sim X_1 + \dots + X_{i-1} + X_{i+1} + \dots \quad (1)$$

$$X_i \sim X_1 + \dots + X_{i-1} + X_{i+1} + \dots \quad (2)$$

Get the following residuals:

$$\mathbf{r}_y = \text{residuals from (1)}$$

$$\mathbf{r}_k^X = \text{residuals from (2)}$$

Plot \mathbf{r}_y vs. \mathbf{r}_k^X : For a valid model, the added-variable plot should produce points randomly scattered around a line through the origin with slope $\hat{\beta}_k$. This is also a useful plot to detect *high influential* data points.

5.5.2 Remedial measure: Linearizing Transformations

1. $\log(Y)$ vs. $\log(X)$, suitable when $\mathbb{E}(Y) = \alpha X_1^{\beta_1} \dots X_p^{\beta_p}$
2. $\log(Y)$ vs. X , suitable when $\mathbb{E}(Y) = \alpha \exp \sum_j X_j \beta_j$
3. $\frac{1}{Y}$ vs. X , suitable when $\mathbb{E}(Y) = \frac{1}{\alpha + \sum_j X_j \beta_j}$

5.5.3 Remedial measure: Box-Cox Transformations of Y also works

Chapter 6 MLR Diagnostics: Collinearity

6.1 Exact Collinearity/ linearly dependent

There exists a set of constants c_1, c_2, \dots, c_p (at least one of them is non-zero) s.t.

$$\sum_{j=1}^p c_j \mathbf{X}_{\cdot j} = 0$$

then the columns of \mathbf{X} are called *linearly dependent* and there is *exact collinearity*.

6.2 What happens if exact collinearity

1. $(\mathbf{X}^T \mathbf{X})^{-1}$ does not exist.
2. The LS estimate $\hat{\beta}$ is not unique.
3. The corresponding linear model is not identifiable.

6.3 Approximate Collinearity

We generally do not need to worry about exact collinearity (**R** can detect it and fix it automatically), but *approximate collinearity*.

$$\sum_{j=1}^p c_j \mathbf{X}_{\cdot j} \approx 0$$

$$\mathbf{X}_{\cdot k} \approx - \sum_{j \neq k} c_j \mathbf{X}_{\cdot j} / c_k$$

A simple diagnostic for this is to obtain the regression of $\mathbf{X}_{\cdot k}$ on the remaining predictors, and if the corresponding R_k^2 is close to 1, we would diagnose approximate collinearity.

$$\mathbf{X}_{\cdot k} \sim \mathbf{X}_{\cdot 1} + \dots + \mathbf{X}_{\cdot k-1} + \mathbf{X}_{\cdot k+1} + \dots \Rightarrow R_k^2$$

6.4 What happens if approximate collinearity: based on $\left(\frac{1}{1-R_k^2} \right)$ (k -th variance inflation factor (VIF), $VIF > 10$)

In a multiple regression $Y = \beta_0 + \beta_1 X_1 + \dots + \beta_p X_p + e$, the LS estimate $\hat{\beta}_k$ is unbiased with variance:

$$\text{Var}(\hat{\beta}_k) = \sigma^2 \left(\frac{1}{1 - R_k^2} \right) \left(\frac{1}{\sum_{i=1}^n (x_{ik} - \bar{x}_{\cdot k})^2} \right)$$

where R_k^2 is the R-square from the regression of $\mathbf{X}_{\cdot k}$ on the remaining predictors. When R_k^2 is close to 1, the variance of $\hat{\beta}_k$ is large.

Consequently we will have:

1. large Mean Square Error
2. large (inflated) p -value to the corresponding t -test, i.e, we could miss a significant predictor.

The quantity $\left(\frac{1}{1-R_k^2}\right)$ is called the k -th variance inflation factor (VIF).

$VIF > 10$ infers collinearity

6.5 Possible symptoms of collinearity

1. high pair-wise (sample) correlation between predictors
2. high VIF
3. high condition number
4. R^2 is relatively large but none of the predictor is significant.

6.6 Global Measure of Collinearity: *condition number of $\mathbf{X}^T \mathbf{X}$*

Condition number of $\mathbf{X}^T \mathbf{X}$:

$$\kappa = (\text{largest eigenvalue/smallest eigenvalue})^{1/2}$$

An empirical rule for declaring collinearity is $\kappa \geq 30$

Note that κ is *not scale-invariant*, so we should *standardize* each column of X (i.e. each column should have *zero mean* and *sample variance* equal to 1, before calculating the condition number).

6.7 What to do with collinearity

1. Remove some predictors from highly correlated groups of predictors.
2. Regularize the model using penalized Least Squares estimation.

Chapter 7 Generalized Least Squares (GLS)

What do we do if the errors are **correlated** or **heteroscedastic**?

Suppose $\varepsilon \sim \mathcal{N}_n(0, \Sigma)$, where Σ is the variance-covariance matrix.

7.1 GLS, Σ known ($\hat{\beta} = (\mathbf{X}^\top \Sigma^{-1} \mathbf{X})^{-1} \mathbf{X}^\top \Sigma^{-1} \mathbf{y}$,

$$RSS = (\mathbf{y} - \mathbf{X}\beta)^\top \Sigma^{-1} (\mathbf{y} - \mathbf{X}\beta)$$

7.1.1 Method 1: Transform back to OLS

$$\mathbf{y} = \mathbf{X}\beta + \varepsilon$$

where $\varepsilon \sim \mathcal{N}_n(0, \Sigma)$ and Σ is a known, symmetric, positive definite covariance matrix.

When the errors are heteroscedastic or correlated:

Transform this problem back to Ordinary Least-Squares (OLS):

1. Assume S^{-1} exists and write

$$\Sigma = SS^\top$$

(We could use, for example, the Cholesky decomposition from linear algebra to obtain S .)

2. Multiply the model equation by S^{-1} on both sides:

$$\mathbf{y} = \mathbf{X}\beta + \varepsilon$$

$$S^{-1}\mathbf{y} = S^{-1}(\mathbf{X}\beta + \varepsilon)$$

$$\underbrace{S^{-1}\mathbf{y}}_{:=\mathbf{y}^*} = \underbrace{S^{-1}\mathbf{X}\beta}_{:=\mathbf{x}^*} + \underbrace{S^{-1}\varepsilon}_{:=\varepsilon^*}$$

$$\mathbf{y}^* = \mathbf{X}^*\beta + \varepsilon^*$$

This implies that

$$\varepsilon^* \sim \mathcal{N}(S^{-1}\mathbf{0}, \underbrace{S^{-1}\Sigma(S^{-1})^\top}_{=\text{Identity}}) = \mathcal{N}(\mathbf{0}, \mathbf{I})$$

since $S^{-1}\Sigma(S^{-1})^\top = S^{-1}SS^\top(S^{-1})^\top = I$

3. For the transformed model, we can solve for β using OLS:

$$\mathbf{y}^* = \mathbf{X}^* \beta + \varepsilon^*$$

where $\mathbf{y}^* = S^{-1} \mathbf{y}$, $\mathbf{X}^* = S^{-1} \mathbf{X}$

So, the estimator for β computes as

$$\begin{aligned}\hat{\beta} &= (\mathbf{X}^{*\top} \mathbf{X}^*)^{-1} \mathbf{X}^{*\top} \mathbf{y}^* \\ &= (\mathbf{X}^\top \underbrace{(S^{-1})^\top}_{=\Sigma^{-1}} S^{-1} \mathbf{X})^{-1} \mathbf{X}^\top \underbrace{(S^{-1})^\top}_{=\Sigma^{-1}} S^{-1} \mathbf{y} \\ &= (\mathbf{X}^\top \Sigma^{-1} \mathbf{X})^{-1} \mathbf{X}^\top \Sigma^{-1} \mathbf{y}\end{aligned}$$

Note that the solution we obtained minimizes:

$$\|\mathbf{y}^* - \mathbf{X}^* \beta\|^2 = (\mathbf{y} - \mathbf{X}\beta)^\top \Sigma^{-1} (\mathbf{y} - \mathbf{X}\beta)$$

7.1.2 Weighted Least Squares (WLS)

Suppose that Σ is a diagonal matrix of unequal error variances:

$$\Sigma = \text{diag}(\sigma_1^2, \sigma_2^2, \dots, \sigma_n^2)$$

The GLS estimate of β minimizes:

$$(\mathbf{y} - \mathbf{X}\beta)^\top \Sigma^{-1} (\mathbf{y} - \mathbf{X}\beta) = \sum_{i=1}^n \frac{(y_i - \mathbf{x}_i^\top \beta)^2}{\sigma_i^2}$$

This problem is known as the Weighted Least-Squares (WLS).

Note that the errors are weighted by

$$w_i = \frac{1}{\sigma_i^2}$$

smaller weights for samples with larger variances.

7.1.3 WLS special example : Replicated Observations

Suppose we collected multiple observations for each \mathbf{x}_i . We use double subscripts to indicate the replicate observations:

$$(\mathbf{x}_i, y_{i1}, y_{i2}, \dots, y_{in_i})$$

Let y_i denote the average of the n_i observations sharing \mathbf{x}_i . Then the residual sum of squares for β equals

$$\sum_{i=1}^n \sum_{j=1}^{n_i} (y_{ij} - \mathbf{x}_i^\top \beta)^2 = \sum_{i=1}^n n_i (y_i - \mathbf{x}_i^\top \beta)^2 + \sum_{i=1}^n \sum_{j=1}^{n_i} (y_{ij} - y_i)^2$$

Minimizing the RSS to solve for β is the same as minimizing the first term on the right only.

$$\hat{\beta} = \underset{\beta}{\operatorname{argmin}} \sum_{i=1}^n n_i(y_i - \mathbf{x}_i^T \beta)^2$$

7.1.4 Method 2: Likelihood Estimation

Model: $\mathbf{y} \sim N_n(\mathbf{X}\beta, \Sigma)$

Log-likelihood:

$$\begin{aligned} & \log(p(\mathbf{y} | \beta, \Sigma)) \\ &= \log \left\{ \frac{1}{(2\pi)^{n/2} |\Sigma|^{1/2}} \exp \left[-\frac{1}{2} (\mathbf{y} - \mathbf{X}\beta)^T \Sigma^{-1} (\mathbf{y} - \mathbf{X}\beta) \right] \right\} \\ &= -\frac{1}{2} (\mathbf{y} - \mathbf{X}\beta)^T \Sigma^{-1} (\mathbf{y} - \mathbf{X}\beta) + \text{Constant}. \end{aligned}$$

Therefore the MLE is given by

$$\hat{\beta}_{mle} = \underset{\beta}{\operatorname{argmin}} (\mathbf{y} - \mathbf{X}\beta)^T \Sigma^{-1} (\mathbf{y} - \mathbf{X}\beta)$$

7.2 GLS, Σ unknown

7.2.1 Method 1: Estimation of Variance (r_i^2)/Standard Deviation Function ($|r_i|$)

$$\sigma_i^2 = \mathbb{E}(\varepsilon_i^2) - (\mathbb{E}(\varepsilon_i))^2$$

Since we assume $E(\varepsilon_i) = 0$, we have

$$\sigma_i^2 = \mathbb{E}(\varepsilon_i^2)$$

Which implies r_i^2 is an estimator of σ_i^2 ; $|r_i|$ is an estimator of the standard deviation σ_i .

Estimate Variance Function $\hat{v}_i(x)$

1. Fit a regression model using OLS, and obtain the residuals r_i .
2. Regress the squared residuals r_i^2 against the appropriate predictor variables.

Denote \hat{v}_i be the fitted value from variance function

$$w_i = \frac{1}{\hat{v}_i}$$

Estimate Standard Deviation Function $\hat{s}_i(x)$

1. Fit a regression model using OLS, and obtain the residuals r_i .
2. Regress the absolute residuals $|r_i|$ against the appropriate predictor variables.

Denote \hat{s}_i be the fitted value from standard deviation function

$$w_i = \frac{1}{(\hat{s}_i)^2}$$

The estimated variances are then placed in the variance-covariance matrix Σ and the regression coefficients are estimated using the WLS (Weighted Least Squares method).

7.2.2 Method 2: iterative approach

1. Start with some initial guess of Σ
2. Use Σ to estimate β
3. Use residuals (since we have known β) to estimate Σ
4. Iterate until convergence.

It looks like a good idea; however the methods will not work if we do not assume some structure about Σ (too many parameters to be estimated).

Based on the application, we can assume a particular structure for Σ that does not involve too many parameters.

Then, we can model β and Σ simultaneously.

For example , for AR(1) times series (auto-regressive model of order 1), the structure of Σ would be:

$$\Sigma = \sigma^2 \begin{pmatrix} 1 & \rho & \rho^2 & \rho^3 & \dots \\ \rho & 1 & \rho & \rho^2 & \dots \\ \dots & \dots & \dots & \dots & \dots \\ \rho^{n-1} & \rho^{n-2} & \dots & \dots & 1 \end{pmatrix}$$

Σ as a function of ρ and σ^2 .

Chapter 8 GLS Diagnostics: Lack of Fit Tests

8.1 Gaussian Assumption

Gaussian Assumption, which can be summarized concisely as:

$$y \sim \mathcal{N}_n(\mathbf{X}\beta, \sigma^2 \mathbf{I})$$

Under these assumptions:

$$\hat{\beta} = (\mathbf{X}^T \mathbf{X})^{-1} \mathbf{X}^T y \sim \mathcal{N}_p(\beta, \sigma^2 (\mathbf{X}^T \mathbf{X})^{-1}),$$

$$\hat{y} = \mathbf{X}\hat{\beta} \sim \mathcal{N}_n(\mathbf{X}\beta, \sigma^2 \mathbf{H})$$

independently,

$$\hat{\sigma}^2 = \frac{RSS}{n-p} = \frac{\|y - \hat{y}\|^2}{n-p} \sim \sigma^2 \frac{\chi_{n-p}^2}{n-p}$$

8.2 When σ^2 is known

Intuition:

If the model is correct, then $\hat{\sigma}^2$ is an unbiased estimate of σ^2 .

If we know σ^2 , we could construct a test based on the ratio $\frac{\hat{\sigma}^2}{\sigma^2}$, a measure of *lack-of-fit*.

In this case we want to test the hypothesis:

$$\begin{cases} H_0 : \text{There is no lack of fit.} \\ H_\alpha : \text{There is lack of fit.} \end{cases}$$

We use the test statistic:

$$\frac{\hat{\sigma}^2}{\sigma^2} = \frac{RSS/(n-p)}{\sigma^2} \sim \frac{\chi_{n-p}^2}{n-p}$$

Lack of fit means the error variance is large related to the value of σ^2 , i.e., the test statistic is large.

Conclude that there is lack of fit (i.e. Reject H_0), if:

$$(n-p) \frac{\hat{\sigma}^2}{\sigma^2} \geq \chi_{n-p}^2(1-\alpha)$$

8.3 When σ^2 is unknown

8.3.1 Hypothesis

If σ^2 is unknown, a general approach is to compare an estimate of σ^2 based on a much bigger/general model.

If we can derive the distribution (under H_0) of $\hat{\sigma}_{\text{LinearModel}}^2 / \hat{\sigma}_{\text{BigModel}}^2$, then we reduce this problem to a two model comparison test problem.

The null hypothesis is the current model:

$$H_0 : \mathbb{E}(y_i) = \mathbf{x}_i^\top \beta, \quad i = 1, 2, \dots, n, \quad \text{for some vector } \beta$$

The more general model is assumed under the alternative hypothesis:

$$H_\alpha : \mathbb{E}(y_i) = f(\mathbf{x}_i), \quad i = 1, 2, \dots, n, \quad \text{for some function } f$$

8.3.2 Under the null hypothesis H_0

$$y_{ij} = \mathbf{x}_i^\top \beta + \varepsilon_{ij}, \text{ some } \beta, \varepsilon_{ij} \sim \text{iid } \mathcal{N}(0, \sigma^2)$$

RSS_0 with $df = n - p$

8.3.3 Under the alternative big-model hypothesis H_α :

$$y_{ij} = f(\mathbf{x}_i) + \varepsilon_{ij}, \text{ some function } f, \varepsilon_{ij} \sim \text{iid } \mathcal{N}(0, \sigma^2)$$

Can we estimate σ^2 for the big model in H_α ?

- The answer is yes, if there is some replication in the data, i.e., there are multiple observations (replicates) for some (at least) of the same \mathbf{x}_i values.
- Schematically we can represent these replicates as:

$$(\mathbf{x}_i, y_{i1}, y_{i2}, \dots, y_{in_i}), \quad i = 1 : m, \quad n = \sum_i n_i$$

RSS_a with $df = n - m = \sum_i (n_i - 1)$, where

$$RSS_a = \sum_{i=1}^m \sum_{j=1}^{n_i} (y_{ij} - \bar{y}_{i.})^2$$

8.3.4 F-Test

All of the degrees of freedom for RSS_a come from the replications. Therefore, with replication we can do an F test for lack of fit:

$$F = \frac{(RSS_0 - RSS_a) / (m - p)}{RSS_a / (n - m)} \sim F_{m-p, n-m}$$

Chapter 9 Polynomials Regression

9.1 Basic Function

$$\begin{aligned}y_i &= f(x_i) + \varepsilon_i \\y_i &= \beta_0 + \sum_{j=1}^d b_j(x_i)\beta_j + \varepsilon_i \\y_i &= \beta_0 + \sum_{j=1}^d \beta_j x_i^j + \varepsilon_i\end{aligned}$$

d is the **degree of the polynomial component**.

9.2 Choose Order d

1. Forward Approach: Keep adding terms until the last added term is not significant.

2. Backward Approach: Start with a large d , and keep eliminating the terms that are not statistically significant, starting with the highest order term.

Once we pick up a d , we do not test the significance of the lower-order terms. (include all the lower-order terms in our model by default)

Reasoning: we do not want our results to be affected by a change of location/scale of the data. $(z_i - 2)^2 = z_i^2 - 4z_i + 4$.

Exception: particular polynomial function (physics law).

9.3 Orthogonal Polynomials

Successive predictors x^j are **highly correlated** introducing multicollinearity problems.

$$y_i = \beta_0 + \beta_1 \mathbf{z}_i + \dots + \beta_d \mathbf{z}_d + \varepsilon_i$$

where $\mathbf{z}_j = a_1 + b_2 x + \dots + \kappa_j x^j$ is a polynomial of order j with coefficients chosen such that $\mathbf{z}_j^T \mathbf{z}_j = 0$

9.4 Piece-wise Polynomials

If the true mean of $E(Y|X = x) = f(x)$ is too wiggly, we might need to fit a higher order polynomial, which is not always a good idea.

Instead we will consider **piece-wise polynomials**:

- 1.we divide the range of x into several intervals, and
- 2.within each interval $f(x)$ is a low-order polynomial, e.g., cubic or quadratic, but the polynomial coefficients change from interval to interval;
- 3.in addition we require the overall $f(x)$ to be continuous up to certain derivatives.

9.5 Cubic Splines

9.5.1 Why Spline

Polynomials: smooth, but each point affects the fit globally.

Piece-wise Polynomials: localizes the influence of each data point, but are not smooth enough.

Splines: combines the beneficail aspects of both approaches.

9.5.2 Settings

A **Cubic Spline** is a curve constructed from sections of cubic polynomials, joined together so that the curve is *continuous up to second derivative*.

The points at which the sections join are called the **knots** of the spline.

We want to define a cubic spline function in the interval $[a, b]$

- Define m knots such that: $a < \xi_1 < \xi_2 < \dots < \xi_m < b$
- A function g defined on $[a, b]$ is a **cubic spline with respect to knots** $\{\xi_i\}_{i=1}^m$ if:
 1. g is a cubic polynomial in each of the $m + 1$ intervals,

$$g(x) = d_i x^3 + c_i x^2 + b_i x + a_i, \quad x \in [\xi_i, \xi_{i+1}]$$

where $i = 0, \dots, m$, $\xi_0 = a$ and $\xi_{m+1} = b$

2. g is continuous up to the 2 nd derivative: since g is continuous up to the 2nd derivative for any point inside

an interval, it suffices to check the following conditions:

$$g^{(0,1,2)}(\xi_i^+) = g^{(0,1,2)}(\xi_i^-), \quad i = 1 : m$$

This expression indicates that the function and the first and second order derivatives are continuous at the knots.

9.5.3 Number of free parameters: $m + 4$

How many free parameters do we need to represent a cubic spline?

- (i) 4 parameters (d_i, c_i, b_i, a_i) for each of the $(m + 1)$ intervals.
- (ii) 3 constraints at each of the m knots (continuity constraints).

The **total number of free parameters** (similar to the number of degrees of freedom) is:

$$4(m + 1) - 3m = m + 4$$

9.5.4 Properties: linear combination also cubic spline

Given knots $\{\xi_i\}_{i=1}^m$, the linear combinations of cubic splines are also cubic splines.

That is, for a set of given knots, the corresponding cubic splines form a linear space (of functions) with $\dim(m + 4)$.

9.5.5 Cubic Splines Basis

A set of basis functions for cubic splines (w.r.t knots $\{\xi_i\}_{i=1}^m$) is given by:

$$h_0(x) = 1$$

$$h_1(x) = x$$

$$h_2(x) = x^2$$

$$h_3(x) = x^3$$

$$h_{i+3}(x) = (x - \xi_i)_+^3, \quad i = 1, 2, \dots, m$$

That is, any cubic spline can be uniquely expressed as:

$$\beta_0 + \sum_{j=1}^{m+3} \beta_j h_j(x)$$

Given knot locations, there are many alternative, but equivalent ways of writing down a basis for cubic splines.

Example 9.1 Other Basis For example, another basis for cubic splines can be the following:

$$h_0(x) = 1$$

$$h_1(x) = x$$

$$h_{i+1}(x) = R(x, \xi_i^*), i = 1, \dots, q-1$$

where

$$\begin{aligned} R(x, z) &= [(z - 1/2)^2 - 1/12] [(x - 1/2)^2 - 1/12] / 4 \\ &\quad - [(|x - z| - 1/2)^4 - 1/2(|x - z| - 1/2)^2 + 7/240] / 24 \end{aligned}$$

9.6 Natural Cubic Splines (NCS)

A cubic spline on $[a, b]$ is a **Natural Cubic Spline** if its *second and third derivatives are zero at a and b* .

9.6.1 Degree of Freedom (Number of free parameters): m

This condition implies that NCS is a linear function in the two extreme intervals $[a, \xi_1]$ and $[\xi_m, b]$. The linear functions in the two extreme intervals are completely determined by their neighboring intervals.

The degree of freedom of NCS's with m knots is:

$$4(m+1) - 3m - 4 = m$$

(We have 4 additional constraints.)

9.6.2 NCS Basis

A Natural Cubic Spline with m knots is represented by m basis functions, for example, one such basis is given by

$$N_1(x) = 1$$

$$N_2(x) = x$$

$$N_{k+2}(x) = d_k(x) - d_{k-1}(x)$$

where

$$d_k(x) = \frac{(x - \xi_k)_+^3 - (x - \xi_m)_+^3}{\xi_m - \xi_k}$$

Each of these derivatives can be seen to have zero second and third derivative for $x \geq \xi_m$.

9.6.3 Note: Waste of Data points

Recall that the **linear functions** in the two extreme intervals are completely determined by the other cubic splines. So data points in the two extreme intervals (i.e., outside the two boundary knots) are wasted since they do not affect the fitting.

9.7 Regression Splines

We can represent the model on the observed n data points using matrix notation:

$$\begin{pmatrix} y_1 \\ y_2 \\ \dots \\ y_n \end{pmatrix}_{n \times 1} = \begin{pmatrix} h_0(x_1) & h_1(x_1) & \dots & h_{p-1}(x_1) \\ h_0(x_2) & h_1(x_2) & \dots & h_{p-1}(x_2) \\ & & \dots & \\ h_0(x_n) & h_1(x_n) & \dots & h_{p-1}(x_n) \end{pmatrix}_{n \times p} \begin{pmatrix} \beta_1 \\ \beta_2 \\ \dots \\ \beta_p \end{pmatrix}_{p \times 1}$$

where our design matrix is the matrix \mathbf{F} of basis functions.

We can find β by solving the problem:

$$\hat{\beta} = \arg \min_{\beta} \|\mathbf{y} - \mathbf{F}\beta\|^2$$

9.8 K -Fold Cross-Validation

How to select the optimal number of knots (or df)?

K-Fold Cross-Validation Steps:

1. Set a fixed number of knots (or df).
2. Divide the set of observations into k groups (or folds).
3. Leave the first fold as a validation set (not used to fit the model). Fit the Regression Spline with a fixed number of knots using the remaining $k - 1$ folds.
4. Calculate the *Mean Square Error* for fold 1: MSE_1 .
5. Repeat the previous steps k times. Each time a new validation set is used to calculate MSE_i .
6. Calculate the average k -fold *Cross-Validation error*:

$$CV(k) = \frac{1}{k} \sum_{i=1}^k MSE_i$$

7. Repeat 2 to 6 with a new number of knots (or df).
8. Select the number of knots that **minimizes** the k -fold CV error or $CV(k)$.

Chapter 10 ANalysis of COVAriance (ANOVA): Basic

These are regression problems where some predictors are quantitative (i.e. numerical) and some are qualitative (i.e. categorical).

10.1 Two level example

For simplicity, we will focus on examples with just two predictors: X (numerical) and D (categorical).

D has two levels: 0 or 1.

General Model

$$y = \beta_0 + \beta_1 x + \beta_2 d + \beta_3(x \cdot d) + \varepsilon$$

Model 1: Coincident regression lines

$$y = \beta_0 + \beta_1 x + \varepsilon$$

Model 1'

$$y = \beta_0 + \beta_2 d + \varepsilon$$

Model 2: Parallel regression lines

$$y = \beta_0 + \beta_1 x + \beta_2 d + \varepsilon$$

Model 3: Regression lines with equal intercepts but different slopes

$$y = \beta_0 + \beta_1 x + \beta_3(x \cdot d) + \varepsilon$$

Model 4: Unrelated regression lines

$$y = \beta_0 + \beta_1 x + \beta_2 d + \beta_3(x \cdot d) + \varepsilon$$

Hierarchical Rule for interactions: *an interaction term will be included in a model only if all its main effects have been included.*

Due to this rule, we would include both β_1 and β_2 , once β_3 is significant. So, we don't consider model 3 this place.

10.2 Two-level Test: t-test

Which model to pick?

1. test whether the interaction term is significant:

$$H_0 : \text{model 2} \quad H_\alpha : \text{model 4}$$

2. if don't reject H_0 , test whether you can further reduce model 2 to model 1

$$H_0 : \text{model 1} \quad H_\alpha : \text{model 2}$$

10.3 Multi-level example

Model the response Y by two predictors X and D , where X is a numerical variable and D is categorical with k levels.

We need to generate $k - 1$ dummy variables: D_2, \dots, D_k where:

$$D_i = \begin{cases} 0, & \text{if not level } i \\ 1, & \text{if level } i \end{cases}$$

Level 1 is the reference level.

Model 0: $Y \sim 1$

Model 1: $Y \sim X$

Model 1': $Y \sim D$

Model 2: $Y \sim D + X$

Model 4: $Y \sim D + X + D : X$

10.4 Multi-level Test: F-test

Note that when D has more than two levels, the difference between model parameter number may not be one, so t-test is no longer appropriate.

- 1) Compare models:

$$H_0 : Y \sim X + D \quad \text{vs.} \quad H_\alpha : Y \sim D + X + D : X$$

If the interaction $D : X$ is significant, stop.

- 2) If X is significant, keep X .

- 2') If D is significant, keep D .

3) If neither X nor D are significant, report the intercept model $Y \sim 1$

Sequential ANOVA We can use the `anova` function to get sequential F-tests. The sequence of F -tests given by `anova(lm(Y ~ X + D + X : D))`

H_0	H_α
$Y \sim 1$	$Y \sim X$
$Y \sim X$	$Y \sim X + D$
$Y \sim X + D$	$Y \sim X + D + X : D$

The sequence of F -tests given by `anova(lm(Y ~ D + X + X : D))`

H_0	H_α
$Y \sim 1$	$Y \sim D$
$Y \sim D$	$Y \sim D + X$
$Y \sim D + X$	$Y \sim D + X + X : D$

Note: Some of the F-stats and p-values from the sequential ANOVA table are **different** from the ones we calculated based on usual F-test (we learned) for comparing two nested models.

Suppose we want to compare:

$$H_0 : Y \sim X \quad \text{vs} \quad H_\alpha : Y \sim X + D$$

The usual F -stat is given by:

$$\frac{(RSS_0 - RSS_a) / (k - 1)}{RSS_a / (n - p_a)} = \frac{(RSS_0 - RSS_a) / (k - 1)}{\hat{\sigma}_a^2}$$

which follows $F_{k-1, n-p_a}$ under the null hypothesis. k is the total number of categories of variable D

The F -stat from the sequential ANOVA table:

$$\frac{(RSS_0 - RSS_A) / (k - 1)}{RSS_A / (n - p_A)} = \frac{(RSS_0 - RSS_A) / (k - 1)}{\hat{\sigma}_A^2}$$

which follows $F_{k-1, n-p_A}$ under the null hypothesis, where RSS_A denotes the RSS from the biggest model $Y \sim X + D + X : D$ and $p_A = 2k$

Chapter 11 Model Adjustment: Variable Selection

11.1 Training and Test Errors

Training data: $(\mathbf{x}_i, y_i)_{i=1}^n$

Test data: $(\mathbf{x}_i, y_i^*)_{i=1}^n$ is an independent (imaginary) data set collected at the same location \mathbf{x}_i 's (also known as in-sample prediction)

Assume the data comes from a linear model:

$\mathbf{y}_{n \times 1}, \mathbf{y}_{n \times 1}^*$ are i.i.d $\sim N_n(\mu, \sigma^2 \mathbf{I}_n)$ and $\mu = \mathbf{X}\beta$

We can also write:

$$\mathbf{y} = \mathbf{X}\beta + \varepsilon$$

$$\mathbf{y}^* = \mathbf{X}\beta + \varepsilon^*$$

with $\varepsilon_{n \times 1}, \varepsilon_{n \times 1}^*$ i.i.d $\sim \mathcal{N}_n(\mathbf{0}, \sigma^2 \mathbf{I}_n)$ are independent.

$$\begin{aligned} \mathbb{E}(\text{Test Error})^2 &= \mathbb{E} \left\| \mathbf{y}^* - \mathbf{X}\hat{\beta} \right\|^2 \\ &= \mathbb{E} \left\| (\mathbf{y}^* - \mathbf{X}\beta) + (\mathbf{X}\beta - \mathbf{X}\hat{\beta}) \right\|^2 \\ &= \mathbb{E} \|\mathbf{y}^* - \mu\|^2 + \mathbb{E} \|\mathbf{X}\beta - \mathbf{X}\hat{\beta}\|^2 \\ &= \mathbb{E} \|\varepsilon^*\|^2 + \text{Tr} \left(\mathbf{X} \text{Cov}(\hat{\beta}) \mathbf{X}^\top \right) \\ &= n \cdot \sigma^2 + \sigma^2 \text{Tr } \mathbf{H} = n \cdot \sigma^2 + p \cdot \sigma^2 \end{aligned}$$

$$\begin{aligned} \mathbb{E}(\text{Train Error})^2 &= \mathbb{E} \|\mathbf{y} - \hat{\mathbf{y}}\|^2 = \mathbb{E} \|(\mathbf{I} - \mathbf{H})\mathbf{y}\|^2 \\ &= \text{Tr} \left((\mathbf{I} - \mathbf{H}) \text{Cov}(\mathbf{y}) (\mathbf{I} - \mathbf{H})^\top \right) \\ &= \sigma^2 \text{Tr}((\mathbf{I} - \mathbf{H})) = (n - p) \cdot \sigma^2 \end{aligned}$$

Index each model (i.e., each subset of the p variables) by a p -dimensional binary vector γ :

$$\gamma = (\gamma_1, \gamma_2, \dots, \gamma_p), \quad \gamma_j = 0/1$$

where $\gamma_j = 1$ indicates that X_j is included in the model, and $\gamma_j = 0$ otherwise.

So there are a total of 2^p possible subsets or sub-models. In particular

$$\gamma = (1, 1, \dots, 1)$$

refers to the full model including all p variables (largest dim), and

$$\gamma = (0, 0, \dots, 0)$$

refers to the intercept-only model (smallest dim).

Suppose that $\mu = \mathbf{X}\beta$ where μ is the mean of \mathbf{y} . If we fit the data \mathbf{y} with respect to model γ , i.e., we fit a linear model with a sub-design matrix X_γ where X_γ contains only columns from X such that $\gamma_j = 1$

We can show that the Testing Error and the Training error for model γ are:

$$\mathbb{E}(\text{Test Error}) = n\sigma^2 + p\sigma^2 + \text{Bias}_\gamma$$

$$\mathbb{E}(\text{Training Error}) = n\sigma^2 - p\sigma^2 + \text{Bias}_\gamma$$

Bigger model (i.e., p large) \rightarrow small Bias, but large variance ($p\sigma^2$) ;

Smaller model (i.e., p small) \rightarrow large Bias, but small variance ($p\sigma^2$).

So to reduce the test error (i.e., prediction error), the key is to find the best trade-off between Bias and Variance.

11.2 Model selection procedures

11.2.1 Testing-based procedures

Testing-based procedures: Select best model based on statistical tests for model comparison.

Backward elimination

- Start with all the predictors in the model.
 - 1. Remove the predictor with highest p -value $> \alpha_0$ (most insignificant).
 - 2. Refit the model, and repeat the above process.
 - 3. Stop when all p -values $\leq \alpha_0$.
- (α_0 is often set to 15% or 20% which is higher than usual)

Forward elimination

- 1. Start with the intercept-only model.
- 2. For all predictors not in the model, check their p -value if being added to the model. Add the one with the lowest p -value $\leq \alpha_0$ (most significant).
- 3. Refit the model, and repeat the above process.
- 4. Stop when no more predictors can be added.

Pros and Cons of Testing-based procedures

- Main advantage: Computation cost is low.
- Due to the "one-at-a-time" nature of adding/dropping variables, this type of procedures does not compare all possible models. So it's possible to miss the "optimal" model.
- It's not clear how to choose α_0 , the cut-off for p -values.

11.2.2 Criterion-based procedures

Criterion-based procedures: Select best model based on an information criteria (combining model fit and model complexity) for model comparison.

1. Score each model according to an information criteria
2. Use a searching algorithm to find the optimal model

Model selection criteria/scores often takes the following form:

$$\text{Training error} + \text{Complexity-penalty}$$

Model Selection Criteria:

AIC/BIC

$$AIC : -2 \times \log lik_{\gamma} + 2p_{\gamma}$$

$$BIC : -2 \times \log lik_{\gamma} + \log(n)p_{\gamma}$$

where p_{γ} is the number of predictors included in model γ

For the linear regression model:

$$-2 \times \log lik_{\gamma} = n \log \frac{RSS_{\gamma}}{n}$$

The lower the AIC/BIC the better. Note that when n is large, adding an additional predictor costs a lot more in BIC than AIC. So AIC tends to pick a bigger model than BIC.

Adjusted $-R^2$ for model γ

$$\begin{aligned} R_a^2 &= 1 - \frac{RSS/(n - p_{\gamma} - 1)}{TSS/(n - 1)} \\ &= 1 - (1 - R^2) \left(\frac{n - 1}{n - p_{\gamma} - 1} \right) \\ &= 1 - \frac{\hat{\sigma}_{\gamma}^2}{\hat{\sigma}_0^2} \end{aligned}$$

The higher the R_a^2 the better.

Mallow's C_p

$$C_p = \frac{RSS_{\gamma}}{\hat{\sigma}^2} + 2p_{\gamma} - n$$

where $\hat{\sigma}^2$ is the estimate of the error variance from the full model. Mallow's C_p behaves very similar to AIC.

Searching Algorithms:

Leap and Bounds:

return *the global optimal solution* among all possible models, but *only feasible for less than 50 variables*.

- Find the p models with the smallest RSS amongst all models of the same size.
- Then evaluate the score on the p models and report the optimal one.

Greedy algorithms:

fast, but *only return a local optimal solution* (which might be good enough in practice).

- Backward: start with the full model and sequentially delete predictors until the score does not improve.
- Forward: start with the null model and sequentially add predictors until the score does not improve.
- Stepwise: consider both deleting and adding one predictor at each stage.

Chapter 12 Model Adjustment: Shrinkage Methods

Find a *trade-off* between *model bias* and *prediction error*.

12.1 Principal Components Regression (PCR)

When we have too many predictors, we need dimensionality reduction in the predictors space. Predictors might be highly correlated.

1. Take matrix \mathbf{X} of predictors and center the columns of \mathbf{X} to have zero mean. Consider \mathbf{X} with no intercept column. (In order to focus on the variation).
2. Find directions of greater variation in the data. (Find the vector that can represent X best.)

12.1.1 Principal Component Analysis (PCA)

The steps to find directions of greater variation in matrix \mathbf{X} :

- Find \mathbf{u}_1 to maximize variance of $\mathbf{u}_1^\top \mathbf{X}$ subject to $\mathbf{u}_1^\top \mathbf{u}_1 = 1$.
- Find \mathbf{u}_2 to maximize variance of $\mathbf{u}_2^\top \mathbf{X}$ subject to $\mathbf{u}_1^\top \mathbf{u}_2 = 0$ and $\mathbf{u}_2^\top \mathbf{u}_2 = 1$
- Continue looking for directions of greatest variation in the data which are orthogonal to the previous ones.
- Continue until the total number of dimensions is exhausted. The principal components are given by the columns of matrix \mathbf{Z} , where

$$\mathbf{Z} = \mathbf{XU}$$

$$[\mathbf{z}_1, \mathbf{z}_2, \dots, \mathbf{z}_m] = [\mathbf{x}_1, \mathbf{x}_2, \dots, \mathbf{x}_m][\mathbf{u}_1, \mathbf{u}_2, \dots, \mathbf{u}_m]$$

\mathbf{z}_i and \mathbf{u}_i are the columns of \mathbf{Z} and \mathbf{U} respectively. \mathbf{U} is called the **rotation matrix**. \mathbf{Z} is a version of the data rotated in such a way that the resulting principal components are orthogonal.

- Each Principal Component is a linear combination of the original variables $\mathbf{x}_1, \mathbf{x}_2, \dots, \mathbf{x}_m$ with weights given by each column \mathbf{u}_i of matrix \mathbf{U} :

$$\mathbf{z}_i = u_{1i}\mathbf{x}_1 + u_{2i}\mathbf{x}_2 + \dots + u_{mi}\mathbf{x}_m$$

- Principal components are very sensitive to outliers.
- The **Mahalanobis distance** can be used to measure the distance of a point to the data mean, after adjusting for correlation in the data.
- Under the multivariate normality assumption in m dimensions, the **Mahalanobis distance**

$$d_i = \sqrt{(\mathbf{x}_i - \mu)^\top \Sigma^{-1} (\mathbf{x}_i - \mu)}$$

can be estimated using the sample estimators for μ and Σ and the quantity d_i^2 follows a χ_m^2 distribution. This can be used to detect outliers in higher dimensions.

12.1.2 After PCA

- Replace model $\mathbf{Y} \sim \mathbf{X}$ by the model $\mathbf{Y} \sim \mathbf{Z}$
- Only need to use the first few columns of \mathbf{Z} as predictors
- Interpretation of the PCAs as predictors might be challenging. We need to use the values of \mathbf{u}_i in the rotation matrix (also called the loadings) for interpretation.
- Sometimes we can make better predictions with a small number of PCs in \mathbf{Z} than with a large number of predictors in \mathbf{X}

12.1.3 Use How many Principal Components?

- The trace of the sample variance-covariance S of Z (total sample variance) is equal to the sum of its eigenvalues:

$$\text{trace}(S) = s_1^2 + s_2^2 + \dots + s_m^2 = \lambda_1 + \lambda_2 + \dots + \lambda_m$$

Since, sample variance-covariance matrix is symmetric, the equation must hold.

- Most of the total variance of a data set is concentrated in the first principal components.

PCs have decreasing explanatory power, so we just need to use the first few:

1. Make a plot of the PCs standard deviations ($\sqrt{\lambda_i}$) vs. the PC index i . This is called the scree plot.
- Look for the PC index i where there is a big change in slope (the elbow) in the scree plot.

Example 12.1

The first comp makes the first point to the second point,... , the fourth comp makes the fourth point to the fifth point, so here we only need the first four comp.

2. Another way is to calculate the cumulative variance explained by the first PCs, and retain the number of PC s explaining between 70% to 90% of the total variation.
3. An alternative way is to discard PCs such that $\lambda_i < \bar{\lambda}$

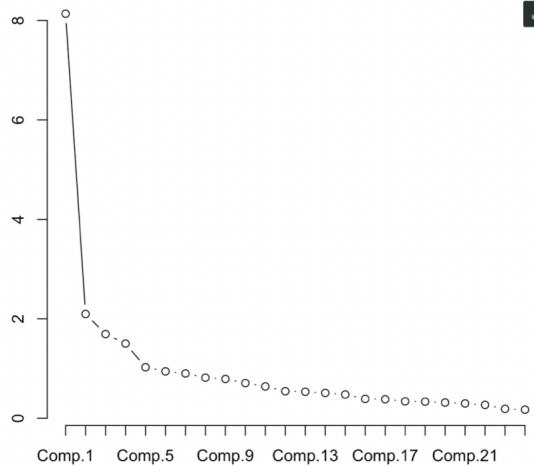


Figure 12.1

12.2 Ridge Regression

- Although the aim of PCR is to reduce dimensionality in the number of predictors, you still have to measure all the predictors since each *PC* is a linear combination of all predictors.
- Ridge regression assumes that after normalization, some of the regression coefficients should not be very large.
- Ridge regression is very useful when you have collinearity and the LS regression coefficients are unstable.
- The method uses a **penalized regression** since the LS minimization problem has a penalty term:

$$\text{minimize} (y - X\beta)^T (y - X\beta) + \lambda \sum_j \beta_j^2$$

for some $\lambda \geq 0$. The penalty term is $\sum_j \beta_j^2$

- Usually predictors are standardized first (centered by their means and scaled by their standard deviations) and the response y is centered.
- The ridge regression estimates are:

$$\hat{\beta} = (\mathbf{X}^T \mathbf{X} + \lambda \mathbf{I})^{-1} \mathbf{X}^T y$$

Or, the β minimize

$$(y - X\beta)^T (y - X\beta) \text{ subject to } \sum_j \beta_j^2 \leq t^2$$

- The parameter λ (or t) should be chosen to have stable estimates of β .
- Note that when $\lambda = 0$ the ridge regression estimation problem reduces to the standard LS problem, while when $\lambda \rightarrow \infty$, $\hat{\beta} \rightarrow 0$
- It is useful to plot the values of $\hat{\beta}_j$ as a function of λ .

- The value of λ can be also chosen using automated methods as **Generalized Cross-Validation (GCV)** (similar to Cross-Validation).
- Ridge regression coefficient estimates are **biased**.

12.3 Lasso Regression

- In this case the estimated $\hat{\beta}$ minimizes:

$$\text{minimize} (y - X\beta)^T (y - X\beta) + \lambda \sum_j |\beta_j|$$

for some $\lambda \geq 0$. The penalty term is $\sum_j |\beta_j|$ (L_1 constraint)

Or, the β minimize

$$(y - X\beta)^T (y - X\beta) \text{ subject to } \sum_j |\beta_j| \leq t$$

- In two-dimensions the constraint defines a square. In higher dimensions it defines a polytope.
- Lasso is useful when the response can be explained by few predictors with zero effect on the remaining predictors (Lasso is similar to a variable selection method).
- When $\beta_j = 0$ the corresponding predictor is eliminated. This is not the case for ridge regression.
- Use Lasso when the effect of predictors is **sparse**. This means that only few predictors will have an effect on the response (e.g. gene expression data) or when number of predictors is large ($p > n$)
- Use the lars R package for Lasso
- Select t in the constraint $\sum_{j=1}^p |\beta_j| \leq t$ by using **Cross-Validation (CV)**
- As t increases, the number of predictors increases.

12.4 Comparing Ridge Regression and Lasso

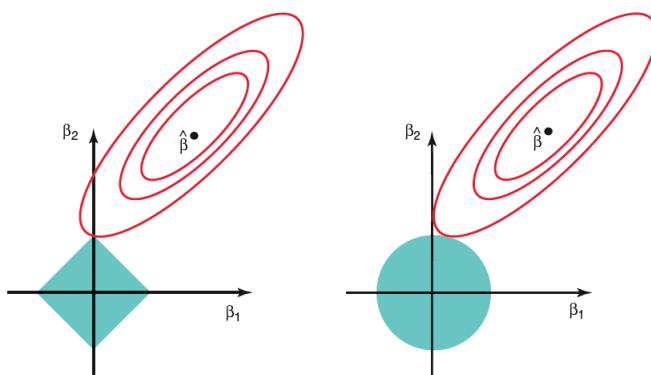


Figure 12.2

In 2 dimensions, Lasso is a square and Ridge is a circle.

- Lasso selects a sub-set of predictors (some coefficients equal to zero).
- Ridge regression performs better when the response is a function of many predictors with coefficients around the same size.
- Lasso will perform better when a relatively small number of predictors have large coefficients and the rest are very small or equal to zero.
- Since the number of predictors is never known a priori, cross-validation can be used to decide which approach is better for a particular data set.

Chapter 13 ANOVA: Comparative Experiments

13.1 Terminology

- **Factor:** an Independent variable. They can be experimental or observational. In our example: Diet
- **Level:** A particular form of the factor. In our example: Levels of the Diet: A, B, C, D
- **Treatments:** Factor levels or factor level combinations (if the study contains more than one factors). They provide insights into mechanisms causing the variation being studied. Control treatments?
- **Complete Randomized Design:** Experimental units are randomly split into r groups, and r treatments are assigned, one per group.

13.2 Data

group 1	$y_{11}, y_{12}, \dots, y_{1n_1}$
group 2	$y_{21}, y_{22}, \dots, y_{2n_2}$
⋮	⋮ ⋮ ⋮ ⋮ ⋮
group r	$y_{r1}, y_{r2}, \dots, y_{rn_r}$

- r is the number of groups
- n_i denotes the number of obs in the i th group
- $n = \sum_{i=1}^r n_i$ is the total sample size
- y_{ij} = observation j for the i th factor.

13.3 ANOVA Model

13.3.1 ANOVA Means Model (Cell Means Model)

$$y_{ij} = \mu_i + \varepsilon_{ij}, i = 1, \dots, r; j = 1, \dots, n_i$$

- y_{ij} : the value of the response in the j th trial for the i th factor.
- μ_i : the population mean for the i th factor level (treatment).
- $\varepsilon_{ij} \sim^{iid} \mathcal{N}(0, \sigma^2)$

13.3.2 Factor Effects Model

Define the effect of factor level i on the response, i.e. the treatment effect as

$$\alpha_i = \mu_i - \mu$$

where μ is the overall mean.

Factor Effects Model:

$$y_{ij} = \mu + \alpha_i + \varepsilon_{ij}, i = 1, \dots, r; j = 1, \dots, n_i$$

$$\varepsilon_{ij} \sim^{iid} \mathcal{N}(0, \sigma^2)$$

- The factor effects model has $r + 1$ model parameters, i.e.

$$(\mu, \alpha_1, \dots, \alpha_r)$$

- In order for the α 's to be (uniquely) estimated, we need to impose restrictions.

- The restrictions on the α 's depend on how μ is defined.

Model	μ Definition	α 's Restriction
Reference Cell	$\mu = \mu_1 \quad \alpha_1 = 0$	
Sum-to-Zero	$\mu = \frac{1}{r} \sum_i \mu_i$	$\sum_i \alpha_i = 0$
Weighted Sum-to-Zero	$\mu = \frac{1}{n} \sum_i n_i \mu_i$	$\sum_i n_i \alpha_i = 0$

- The default in R is the Reference Cell model.

13.4 Model Properties

$$- E(y_{ij}) = \mu_i$$

$$- \text{Var}(y_{ij}) = \text{Var}(\varepsilon_{ij}) = \sigma^2$$

Thus, all observations have the same variance, regardless of factor level.

- $\varepsilon_{ij} \sim \mathcal{N}(0, \sigma^2)$ and independent

- $y_{ij} \sim \mathcal{N}(\mu_i, \sigma^2)$ and independent.

We can re-state the model as

$$y_{ij} \text{ are independent } \mathcal{N}(\mu_i, \sigma^2)$$

13.5 Model Estimation

Minimize the sum of squared deviations of the observations around their expected values with respect to the parameters:

$$Q = \sum_{i=1}^r \sum_{j=1}^{n_i} (y_{ij} - \mathbb{E}(y_{ij}))^2$$

If we re-write Q we have

$$Q = \sum_j (y_{1j} - \mu_1)^2 + \sum_j (y_{2j} - \mu_2)^2 + \dots + \sum_j (y_{rj} - \mu_r)^2$$

So the **least squares estimator** of μ_i , denoted by $\hat{\mu}_i$ is

$$\hat{\mu}_i = \bar{y}_{i\cdot} = \frac{1}{n_i} \sum_{j=1}^{n_i} y_{ij}$$

Using the appropriate constraints, we can easily extract the estimators for μ and α_i .

- The *LS* fit for y_{ij} is the corresponding group mean

$$\hat{y}_{ij} = \bar{y}_i$$

- Residuals

$$r_{ij} = y_{ij} - \hat{y}_{ij} = y_{ij} - \bar{y}_i$$

- RSS

$$\sum_{i=1}^r \sum_{j=1}^{n_i} (y_{ij} - \bar{y}_{i\cdot})^2$$

i.e. the within-group variation.

Source of Variation	SS	df	MS
Between Groups	$FSS = \sum n_i (\bar{y}_{i\cdot} - \bar{y}_{..})^2$	$r - 1$	$\frac{FSS}{r-1}$
Error (within Groups)	$RSS = \sum \sum (y_{ij} - \bar{y}_{i\cdot})^2$	$n - r$	$\frac{RSS}{n-r}$
Total	$TSS = \sum \sum (y_{ij} - \bar{y}_{..})^2$	$n - 1$	

13.6 F-test

- We want to test whether the means of the groups are really different. We can express this as

$$\begin{cases} H_0 : \mu_1 = \mu_2 = \dots = \mu_r \\ H_\alpha : \text{not all } \mu_i, i = 1, \dots, r \text{ are equal} \end{cases}$$

- or in terms of models

$$\begin{cases} H_0 : y_{ij} = \mu + \varepsilon_{ij} \\ H_\alpha : y_{ij} = \mu + \alpha_i + \varepsilon_{ij} \end{cases}$$

- They are two nested models, so we can use the F -test

$$\frac{(RSS_0 - RSS_\alpha) / (r - 1)}{RSS_\alpha / (n - r)} \sim F_{r-1, n-r}$$

under H_0 .

- The test statistic can also be written as

$$\frac{FSS / (r - 1)}{RSS / (n - r)} = \frac{\text{Between-group Variation} / (r - 1)}{\text{Within-group Variation} / (n - r)}$$

where FSS, RSS are defined in the ANOVA table.

13.7 Diagnostics for ANOVA Models

- Check for outliers/ unusual observations.
- Check the residuals vs. fitted values plot for departures from the constant variance assumption.
- Check the Q-Q plot for departures from the normality assumption.

Levene's Test for Equality of Variances:

- Run Regression $\text{abs(residuals)} \sim X$, i.e. use abs(residuals) as the response in a new one-way ANOVA.
- If the p-value for the F-test is **greater** than 1% level, then we conclude that there is no evidence of a non-constant variance.

13.8 Inference for Factor Level Means (function about the μ_i s)

13.8.1 A single factor level mean

- Estimation of

$$\mu_i : \hat{\mu}_i = \bar{y}_i$$

- Distribution of

$$\hat{\mu}_i : E(\hat{\mu}_i) = \mu_i, \quad \text{Var}(\hat{\mu}_i) = \frac{\sigma^2}{n_i}$$

- The estimated variance of \bar{y}_i . is

$$s_{\bar{y}_i}^2 = \frac{1}{n_i} \cdot \frac{RSS}{n - r}$$

- Under the ANOVA model assumptions $\frac{\bar{y}_i - \mu_i}{s_{\bar{y}_i}}$ is distributed as T_{n-r}

- Confidence Interval for μ_i :

$$\mu_i \in \bar{y}_i \pm T_{n-r}(\alpha/2)s_{\bar{y}_i}$$

13.8.2 A difference between two factor level means

The difference between two factor level means (pairwise comparison) is defined as

$$D = \mu_i - \mu_{i'}$$

- Estimation of D :

$$\hat{D} = \bar{y}_i - \bar{y}_{i'}$$

- Distribution of \hat{D} :

$$E(\hat{D}) = \mu_i - \mu_{i'}, \text{Var}(\hat{D}) = \sigma^2 \left(\frac{1}{n_i} + \frac{1}{n_{i'}} \right)$$

The estimated variance of \hat{D} is

$$s_{\hat{D}}^2 = \frac{RSS}{n-r} \cdot \left(\frac{1}{n_i} + \frac{1}{n_{i'}} \right)$$

- Under the ANOVA model assumptions

$$\frac{\hat{D} - D}{s_{\hat{D}}} \text{ is distributed as } T_{n-r}$$

- Confidence Interval for D :

$$D \in \hat{D} \pm T_{n-r}(\alpha/2)s_{\hat{D}}$$

- Hypothesis Test for D :

$$\begin{cases} H_0 : \mu_i = \mu_{i'} \\ H_\alpha : \mu_i \neq \mu_{i'} \end{cases} \Leftrightarrow \begin{cases} H_0 : \mu_i - \mu_{i'} = D = 0 \\ H_\alpha : \mu_i - \mu_{i'} \neq 0 \end{cases}$$

The test statistic is

$$t = \frac{\hat{D}}{s_{\hat{D}}} \sim T_{n-r}$$

13.8.3 A contrast among factor level means

A contrast is a comparison involving two or more level means:

$$L = \sum_{i=1}^r c_i \mu_i, \quad \text{where } \sum_{i=1}^r c_i = 0$$

- Estimation of L :

$$\hat{L} = \sum_{i=1}^r c_i \bar{y}_i$$

- Distribution of \hat{L} :

$$E(\hat{L}) = \sum_{i=1}^r c_i \mu_i, \text{Var}(\hat{L}) = \sigma^2 \sum_{i=1}^r \frac{c_i^2}{n_i}$$

The estimated variance of \hat{L} is

$$s_{\hat{L}}^2 = \frac{RSS}{n-r} \cdot \sum_{i=1}^r \frac{c_i^2}{n_i}$$

- Under the ANOVA model assumptions $\frac{\hat{L}-L}{s_{\hat{L}}}$ is distributed as T_{n-r}

- Confidence Interval for L :

$$L \in \hat{L} \pm T_{n-r}(\alpha/2)s_{\hat{L}}$$

- Hypothesis Testing for L :

$$\begin{cases} H_0 : L = 0 \\ H_\alpha : L \neq 0 \end{cases}$$

The test statistic is

$$t = \frac{\hat{L}}{s_{\hat{L}}} \sim T_{n-r}$$

13.8.4 A linear combination of factor level means

$$L = \sum_{i=1}^r c_i \mu_i, \quad \text{no restrictions on } c'_i s$$

- Point estimator and estimated variance same as before.

- Single Degree of Freedom Tests

$$\begin{cases} H_0 : L = c \\ H_\alpha : L \neq c \end{cases}$$

The test statistic here is

$$F = t^2 = \left(\frac{\hat{L} - c}{s_{\hat{L}}} \right)^2 \sim F_{1,n-r}$$

13.9 Limitations of Inference Procedures

The confidence coefficient $1 - \alpha$ for the estimation procedures described is a statement confidence coefficient and applies **only to a particular estimate, not to a series of estimates**.

Similarly the specified Type I error rate α applies only to a particular test and not to a series of tests.

13.10 Bonferroni Correction $\frac{\alpha}{m}$

Example 13.1 If the confidence coefficients are 95% for all individual μ_i , the confidence coefficient will be $(95\%)^n < 95\%$ for family $f(\mu_1, \dots, \mu_n)$.

So the 95% confidence interval of family will be **wider** than individual.

When? The family of interest is a particular set of pairwise comparisons, contrasts, or linear combinations that is specified by the user.

- Suppose m is the number of statements in the family.
- In order to control the family wise error rate to be α , we need to reduce the error rate for each individual comparison to be α/m .
- That is we need to increase the significance level from $(1 - \alpha)$ to $(1 - \alpha/m)$
- Not applicable when m is large, since the CIs would be too wide due to the increase of the significant level.
i.e. We use $1 - \frac{5\%}{n}$ for all μ_i to form 95% $f(\mu_1, \dots, \mu_n)$

13.11 Tukey's Paired Comparison Procedures

When? the family of interest is a set of all pairwise comparisons of factor level means, i.e. it consists of estimates of all pairs $D = \mu_i - \mu_{i'}$

A confidence interval is given by

$$D \in \hat{D} + \frac{q(\alpha/2; r, n - r)}{\sqrt{2}} s(\hat{D})$$

where $q(\alpha/2; r, n - r)$ refers to the $\alpha/2$ upper quantile of the studentized range for r means and $n - r$ degrees of freedom.

The coverage probability is exact when the sample sizes in each group are identical and is approximate otherwise.

Remark: The studentized range refers to the distribution of

$$\max_{i \neq j} \sqrt{n} (\bar{y}_i - \bar{y}_j) / \hat{\sigma}$$

where \bar{y}_i and \bar{y}_j are sample means from independent samples of size n from normal distributions with common means and variance σ^2 .

Note: Tukey is always better than Bonferroni and Scheffe in pairwise comparisons.

13.12 Scheffe's Method for Contrasts

When? The family of interest is the set of contrasts among the factor level means:

$$L = \sum c_i \mu_i, \text{ where } \sum c_i = 0$$

An confidence interval is given by

$$L \in \hat{L} + (r - 1)F_{r-1, n-r}(\alpha)s_{\hat{L}}$$

Chapter 14 Two Way ANOVA

Single-factor:

1. Do not explore the entire space of treatment combinations.
2. Interactions cannot be estimated.
3. Full randomization is not possible.
4. Multiple stages increase complexity of the analysis.

MultiFactor:

1. Efficient replication.
2. Assessment of Interactions.
3. Validity of Findings.

14.1 Factor Effects Model for Two Factors

$$y_{ijk} = \mu + \alpha_i + \beta_j + (\alpha\beta)_{ij} + \varepsilon_{ijk}$$

	Sum	Average
Cell (i, j)	$y_{ij.} = \sum_{k=1}^n y_{ijk}$	$\bar{y}_{ij.} = \frac{y_{ij.}}{n}$
Row i	$y_{i..} = \sum_{j=1}^b \sum_{k=1}^n y_{ijk}$	$\bar{y}_{i..} = \frac{y_{i..}}{bn}$
Column j	$y_{.j} = \sum_{i=1}^a \sum_{k=1}^n y_{ijk}$	$\bar{y}_{.j} = \frac{y_{.j}}{an}$
Overall	$y_{...} = \sum_{i=1}^a \sum_{j=1}^b \sum_{k=1}^n y_{ijk}$	$\bar{y}_{...} = \frac{y_{...}}{nab}$

- Using least squares method, the estimated treatment means are:

$$\hat{\mu}_{ij} = \bar{y}_{ij}$$

- The factor effects estimators depend on the constraints that we impose. For example, under the sum-constraints we have

$$\hat{\alpha}_i = \bar{y}_{i..} - \bar{y}_{...}, \quad \hat{\beta}_j = \bar{y}_{.j} - \bar{y}_{...}$$

$$(\hat{\alpha}\hat{\beta})_{ij} = \bar{y}_{ij} - \bar{y}_{i..} - \bar{y}_{.j} + \bar{y}_{...}$$

- The fitted values and residuals compute as usual as

$$\hat{y}_{ijk} = \bar{y}_{ij}, \quad r_{ij} = y_{ijk} - \hat{y}_{ijk}$$

14.2 Interaction Plots

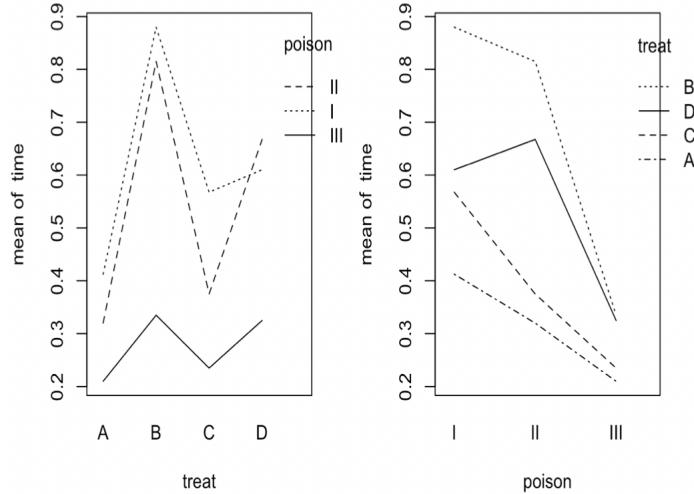


Figure 14.1

If the lines are not parallel, interaction is presented.

14.3 Partitioning of Total Sum of Squares

$$\begin{array}{c}
 \underbrace{y_{ijk} - \bar{y}_{...}}_{\text{Total Deviation}} = \underbrace{\bar{y}_{ij} - \bar{y}_{...}}_{\substack{\text{Deviation of estimated} \\ \text{treatment mean around} \\ \text{overall mean}}} + \underbrace{y_{ijk} - \bar{y}_{ij.}}_{\substack{\text{Deviation} \\ \text{around estimated} \\ \text{treatment mean}}}
 \end{array}$$

$$TSS = FSS + RSS$$

where

$$TSS = \sum_i \sum_j \sum_k (y_{ijk} - \bar{y}_{...})^2$$

$$FSS = n \sum_i \sum_j (\bar{y}_{ij.} - \bar{y}_{...})^2$$

$$RSS = \sum_i \sum_j \sum_k (y_{ijk} - \bar{y}_{ij.})^2 = \sum_i \sum_j \sum_k e_{ijk}^2$$

14.4 Partitioning of Treatment Sum of Squares

$$\underbrace{\bar{y}_{ij.} - \bar{y}_{...}}_{\text{Deviation of estimated treatment mean around overall mean}} = \underbrace{\bar{y}_{i..} - \bar{y}_{...}}_{\text{A main effect}} + \underbrace{\bar{y}_{.j.} - \bar{y}_{...}}_{\text{B main effect}} + \underbrace{\bar{y}_{ij.} - \bar{y}_{i..} - \bar{y}_{.j.} + \bar{y}_{...}}_{\text{A B interaction effect}}$$

$$FSS = SSA + SSB + SSAB \text{ (Orthogonal Decomposition)}$$

where

$$SSA = nb \sum_i (\bar{y}_{i..} - \bar{y}_{...})^2$$

$$SSB = na \sum_j (\bar{y}_{.j.} - \bar{y}_{...})^2$$

$$SSAB = n \sum_i \sum_j (\bar{y}_{ij.} - \bar{y}_{i..} - \bar{y}_{.j.} + \bar{y}_{...})^2$$

14.5 ANOVA Table

Source of Variation	SS	df	MS
Factor A	SSA	$a - 1$	$MSA = \frac{SSA}{a-1}$
Factor B	SSB	$b - 1$	$MSB = \frac{SSB}{b-1}$
AB Interactions	SSAB	$(a - 1)(b - 1)$	$MSAB = \frac{SSAB}{(a-1)(b-1)}$
Error	RSS	$ab(n - 1)$	$MSE = \frac{RSS}{ab(n-1)}$
Total	TSS	$nab - 1$	

14.6 F-test

- In order to test for the statistical significance of the interaction terms, we use partial F -tests. So, we fit a main effects model (i.e. no interactions)

$$y_{ijk} = \mu + \alpha_i + \beta_j + e_{ijk}$$

- Then, we compare the two nested models:

$$\begin{cases} H_0 : \text{smaller model with } p_0 \text{ coefficients} \\ H_\alpha : \text{larger model with } p_\alpha \text{ coefficients} \end{cases}$$

The F -test is formulated as

$$F = \frac{(RSS_0 - RSS_\alpha) / (p_\alpha - p_0)}{MSE_\alpha} \sim F_{p_\alpha - p_0, n - p_\alpha} \text{ under the } H_0$$

We can also perform F-tests directly using the ANOVA table, where for the interaction term we have:

$$F_{AB} = \frac{MSAB}{MSE} \sim F_{(a-1)(b-1), nab-1}$$

Hierarchy principle: we test for main effects only if the interaction term is not statistically significant.

14.7 Estimation of Factor Level Means

When interactions are not statistically significant, we analyze the factor level means:

- Factor Level Means:

$$\hat{\mu}_{i\cdot} = \bar{y}_{i\cdot}, s_{\hat{\mu}_{i\cdot}}^2 = \frac{MSE}{bn}$$

- Differences of Factor Level Means:

$$\hat{\mu}_{i\cdot} - \hat{\mu}_{i'\cdot} = \bar{y}_{i\cdot} - \bar{y}_{i'\cdot}, s_D^2 = \frac{2MSE}{bn}$$

- Contrasts of Factor Level Means:

$$\hat{L} = \sum c_i \hat{\mu}_{i\cdot} = \sum c_i \bar{y}_{i\cdot}, s_{\hat{L}_{(i)}}^2 = \frac{MSE}{bn} \sum c_i^2$$

where $\sum c_i = 0$

For individual hypothesis test and Cls, the multiplier is $T_{(n-1)ab}(\alpha/2)$.

For family hypothesis tests/intervals, we select the desired family multiplier:

- Tukey Multiplier:

$$\frac{1}{\sqrt{2}} q_{a,(n-1)ab}(1 - \alpha)$$

- Bonferroni Multiplier: $B = T_{(n-1)ab}(1 - \alpha/2m)$, where m refers to the number of multiple comparisons.

- Scheffé Multiplier:

- $S^2 = (b-1)F_{b-1,(n-1)ab}(1 - \alpha)$, if the contrasts involve the $\mu_{i\cdot}$ and

- $S^2 = (a-1)F_{a-1,(n-1)ab}(1 - \alpha)$, if the contrasts involve the $\mu_{j\cdot}$.

14.8 Estimation of Treatment Means

When interactions are statistically significant, we analyze the treatment means:

- Treatment Means:

$$\hat{\mu}_{ij} = \bar{y}_{ij}, \quad s_{\hat{\mu}_{ij}}^2 = \frac{MSE}{n}$$

- Differences of Treatment Means:

$$\hat{D} = \hat{\mu}_{ij} - \hat{\mu}_{i'j'} = \bar{y}_{ij} - \bar{y}_{i'j'}, \quad i, j \neq i', j' \text{ and } s_{\hat{D}}^2 = \frac{2MSE}{n}$$

- Contrasts of Treatment Means:

$$\hat{L} = \sum \sum c_{ij} \hat{\mu}_{ij} = \sum \sum c_{ij} \bar{y}_{ij}, \text{ where } \sum \sum c_{ij} = 0$$

$$\text{with variance } s_{\hat{L}}^2 = \frac{MSE}{n} \sum c_{ij}^2$$

For individual hypothesis test and CIs, the multiplier is $T_{(n-1)ab}(\alpha/2)$.

For family hypothesis tests/intervals, we choose the desired family multiplier:

- Tukey Multiplier: $\frac{1}{\sqrt{2}} q_{ab, (n-1)ab}(1 - \alpha)$

- Bonferroni Multiplier: $B = T_{(n-1)ab}(1 - \alpha/2m)$, where m refers to the number of multiple comparisons.

- Scheffé Multiplier: $S^2 = (ab - 1) F_{ab-1, (n-1)ab}(1 - \alpha)$, if the contrasts involve the $\mu_{i..}$.

Chapter 15 Two Way ANOVA: Special Cases

15.1 Unbalanced ANOVA (Use Partial F -Test or ANOVA type III in R)

- When the treatment sample sizes are **unequal**, the analysis of variance for two-factor studies becomes more complex.
- The least-squares equations are no longer of a simple structure and the regular analysis of variance formulas are now inappropriate.
- Furthermore, the factor effect component sum of squares are no longer orthogonal; that is, they **do not sum up to TSS**.

$$y_{ijk} = \mu + \alpha_i + \beta_j + (\alpha\beta)_{ij} + \varepsilon_{ijk}$$

- Due to the lack of orthogonality, the **ANOVA F-tests are not applicable**.

15.1.1 Use Partial F -Test

- We will express the ANOVA model as a regression model with **indicator (dummy) variables**.
- We need $a - 1$ indicator variables for factor A main effects and $b - 1$ indicator variables for factor B main effects. The interactions correspond to the cross products of the indicator variables for A and B.

Use **Partial F -Test**.

15.1.2 ANOVA type III in R

This type tests for the presence of an effect given that both the other effects are in the model.

Example 15.1 Compute SSTotal

$$RSS_0 - RSS_\alpha = (SSA + SSR) - SSR = SSA = 2.1340$$

$$(SSB + SSR) - SSR = SSB = 11.7375$$

$$(SSAB + SSR) - SSR = SSAB = 1.9800$$

$$SST = 2.1340 + 11.7375 + 1.9800 + 7.9151 = 23.7666$$

```
Anova(lm(1/time ~ treat*poison, data=newrats), type="III")

## Anova Table (Type III tests)
##
## Response: 1/time
##           Sum Sq Df F value    Pr(>F)
## (Intercept) 15.0605  1 66.5967 1.298e-09 ***
## treat        2.1340  3  3.1455  0.03723 *
## poison       11.7375  2 25.9514 1.225e-07 ***
## treat:poison  1.9800  6  1.4592  0.22073
## Residuals    7.9151 35
## ---
## Signif. codes: 0 '***' 0.001 '**' 0.01 '*' 0.05 '.' 0.1 ' ' 1
```

Figure 15.1

15.2 Balanced ANOVA with $n = 1$ (Tukey's Test)

- Only one observation in each cell, so we cannot fit the interaction model.
- There are no degrees of freedom left for estimating the error.
- RSS = 0 when the model includes main effects and interaction term. (error is 0)
- All F-tests are valid, but the interaction model is not a candidate model.

15.2.1 Tukey's Test for Additivity

Consider the following model that includes interactions:

$$y_{ij} = \mu + \alpha_i + \beta_j + \theta\alpha_i\beta_j + \varepsilon_{ij}$$

Here, we assume that the interactions are of *multiplicative* nature, i.e.

$$(\alpha\beta)_{ij} = \theta\alpha_i\beta_j$$

- Consider the SSA, SSB as before and:

$$SSAB^* = \frac{\left(\sum_i \sum_j (\bar{y}_i - \bar{y}_{..}) (\bar{y}_{.j} - \bar{y}_{..}) y_{ij} \right)^2}{\sum_i (\bar{y}_{i.} - \bar{y}_{..})^2 \sum_j (\bar{y}_{.j} - \bar{y}_{..})^2}$$

- The TSS is computed as usual and is decomposed as

$$TSS = SSA + SSB + SSAB^* + SSRem^*$$

where the remainder is

$$SSRem^* = TSS - SSA - SSB - SSAB^*$$

- We want to test the following hypothesis

$$\begin{cases} H_0 : \theta = 0 & (\text{no interactions}) \\ H_\alpha : \theta \neq 0 & (\text{interactions}) \end{cases}$$

which is essentially a test for model additivity. - The test statistic computes as

$$F^* = \frac{SSAB^*/1}{SSRem^*/(ab - a - b)}$$

Chapter 16 Experimental Designs: Introduction

16.1 Experimental vs. Observational Study

Experimental Study is a scientific procedure undertaken to make a discovery, test a hypothesis or verify a claim.

Observational Study is one in which the experimenter observes the effect of a factor on the response, or measures an outcome without an attempt to affect the outcome by intervention.

16.2 Principles of Experimental Design

Randomization

- Random allocation of treatment and order
- Ensures that collected data are IID random variables
- Averages-out the effects of exogenous factors.

Replication

- Estimate of experimental error
- Higher Precision

Blocking

- Higher precision when comparisons of factors are made
- Reduced variability transmitted from nuisance factors.

16.3 Randomization Test

- A test based directly on re-randomizing - with the same kind of randomization originally used to assign the treatments - is called a **randomization test**.

- *Advantage:* No need for any distributional assumptions (independence, normality, etc.) - just need to assume that the treatment randomization was performed properly.

- *Disadvantage:* Requires more computation, and you must implement for yourself or use specialized software.

Chapter 17 Experimental Designs: Blocking

17.1 Randomized Complete Block Design (RCBD) Model

Example:

- Treatment Factor ‘variety’: 8 levels
- Block Factor ‘block’: 5 levels
- Observe that we have *only one* observation per treatment-block combination.

```
xtabs(yield~variety + block, oatvar)

##          block
## variety   I   II  III  IV   V
##      1 296 357 340 331 348
##      2 402 390 431 340 320
##      3 437 334 426 320 296
##      4 303 319 310 260 242
##      5 469 405 442 487 394
##      6 345 342 358 300 308
##      7 324 339 357 352 220
##      8 488 374 401 338 320
```

Figure 17.1

Suppose that there are r treatments (factor levels) and n_b blocks.

$$y_{ij} = \mu + \tau_i + \beta_j + \varepsilon_{ij}$$

where

- μ . is a constant
- τ_i are the treatment effects
- β_j are the block effects
- ε_{ij} are independent $\mathcal{N}(0, \sigma^2)$
- $i = 1, \dots, n_b$ (total number of blocks), $j = 1, \dots, r$ (total number of treatments)

Remarks

- y_{ij} is the response for the j th treatment in the i th block.

- There is a single observation per block. This implies that we have a limited ability to detect an interaction between treatment and block. So, we are working with the additive model.
- We can check for treatment and block main effects, but blocking is a feature of the design which means that if insignificant, we cannot gain the degrees of freedom.

ANOVA Display for the RCBD (Two factors: Treatments and Blocks)

Source of Variation	Sum of Squares	Degrees of Freedom	Mean Square	F_0
Treatments	$SS_{\text{Treatments}}$	$r - 1$	$\frac{SS_{\text{Treatments}}}{r-1}$	$\frac{MS_{\text{Treatments}}}{MS_g}$
Blocks	SS_{Blocks}	$n_b - 1$	$\frac{SS_{\text{Blocks}}}{n_b-1}$	
Error	SS_E	$(r - 1)(n_b - 1)$	$\frac{SS_g}{(r-1)(n_b-1)}$	
Total	SS_T	$rn_b - 1$		

17.2 Latin Squares

17.2.1 Example

- A, B, C: 3 treatments
- Day of week (Monday, Wednesday, Friday): Blocking Variable
- Operator ID: 1, 2, 3: Blocking Variable

		Operator		
		Day	1	2
		Monday	B	A
		Wednesday	A	C
		Friday	C	B

* Each operator runs each treatment, and all treatments are run on each day

17.2.2 Features of a Latin Square Design

- There are r treatments.
- There are 2 blocking variables, each containing r classes.
- Each row and each column in the design square contains all treatments.
- Each treatment is assigned to each block only once.

Advantages

- Reduces more experimental error than with 1 blocking factor.
- Small scale studies can isolate important treatment features.
- Repeated measures designs can remove order effects.

Disadvantages

- Each blocking factor must have r levels.
- No interactions among factors.
- With small r , we have very few error degrees of freedom.
- Complex Randomization.

17.2.3 Randomization in Latin Square Designs

- Determine r , the number of treatments, row blocks, and column blocks.
- Select a Standard Latin Square (from tables or with software).
- Use Capital Letters to represent treatments (A, B, C, \dots) and randomly assign treatments to labels.
- Randomly assign Row Block levels to Square Rows.
- Randomly assign Column Block levels to Square Columns.

17.2.4 Latin Square Model

$$y_{ijk} = \mu + \tau_i + \beta_j + \gamma_k + e_{ijk}$$

where

- μ is a constant
- τ_i treatment effect (latin letter)
- β_j (column) blocking effect
- γ_k (row) blocking effect
- e_{ijk} are independent $\mathcal{N}(0, \sigma^2)$
- $i, k, j = 1, \dots, r$

NOVA Display for the Latin Square Model (Three factors: Treatments, Rows and Cols)

AOV	df
Rows (blocks)	$r - 1$
Cols (blocks)	$r - 1$
Treatments	$r - 1$
Error	$(r - 1)(r - 2)$
Total	$(r^2 - 1)$

17.3 Balanced Incomplete Block Design (BIBD)

- [→] Why Balanced?

Each pair of treatments occur together λ times.

- [→] Why Incomplete?

Cannot fit all treatments in each block.

Notation

- t treatments

- b blocks

- k treatments per block (block size)

- r times each treatment occurs

- $N = t \cdot r = b \cdot k$ observations in total

- Treatment i occurs in r blocks.

- To have balance, each other treatment is equally likely to be treatment i in a block.

- Since there are $k - 1$ other units in a block and $t - 1$ other treatments, the number of times each pair of treatments appears in the same block is

λ = the probability that there existst treatment j in remaining units in one block \times the number of blocks that have treatment

$$= \frac{k-1}{t-1} r = \frac{r(k-1)}{t-1}$$

where λ is an integer.

Examples

- $t = 3, b = 3, k = 2 \rightarrow r = 2, \lambda = 1$.

$t = \# \text{ treatments } \{A, B, C\} = 3$

The form of square: $k \times b = 2 \times 3$.

$r = \# A \text{ occurs} = \# B \text{ occurs} = \# C \text{ occurs} = 2$

$\lambda = \# A \text{ and } B \text{ in one block} = \# A \text{ and } C \text{ in one block} = \# B \text{ and } C \text{ in one block} = 1$

Block		
1	2	3
A	B	A
B	C	C

- $t = 4, k = 2, b = 6 \rightarrow r = 3, \lambda = 1$

Block					
1	2	3	4	5	6
A	A	A	B	B	C
B	C	D	C	D	D

17.4 BIBD Remarks

Advantages:

- A BIBD enables us to run an experiment when the size of the available blocks of experimental units is smaller than the number of treatments.
- Estimates of treatment effects have equal precision and expressions for the variances of the estimated cell means and of contrasts of treatment means or effects are relatively simple.
- The presence of balance permits the use of Scheffé and Tukey procedures for the analysis of treatment effects.

Disadvantages:

- BIBD exist only for certain combinations of numbers of treatments, block sizes, and numbers of blocks.
- The assumption that there are no interactions between the blocking variable and the treatments is restrictive.
- The analysis of a BIBD is more complex than that of a RCBD.

$$y_{ij} = \mu + \tau_i + \beta_j + e_{ij}$$

- μ constant

- τ_i treatment effects

- β_j the block effects
- e_{ij} independent $N(0, \sigma^2)$

Remarks

- Not all y_{ij} exist, because of incompleteness.
- Non-orthogonality of treatments and blocks.

Chapter 18 Linear Models with Random Effects

18.1 Random Effects

One-way ANOVA model

$$y_{ij} = \mu + \alpha_i + \varepsilon_{ij}$$

Previously, we assumed the α_i s were parameters: fixed, unknown values. (We also gave a restriction, to make them identifiable.) These are fixed effects, corresponding to a fixed factor.

Now suppose that the α_i s are unobserved random variables:

$$\alpha_i \sim N(0, \sigma_\alpha^2)$$

and assume they are independent of each other and of the ε_{ij} s.

These α_i s are called random effects, and the corresponding factor variable is a random factor.

- **Fixed effects** are appropriate when the levels of the factor are individually important or meaningful (e.g. treatments in a designed experiment, levels of education).
- **Random effects** are appropriate when the levels of the factor are meaningful only as representatives of a more general collection (e.g. as if sampled from a population, or representative of some hypothetical population).

- For the one-way ANOVA model with a random factor, the random effects satisfy

$$E(\alpha_i) = 0 \quad \text{Var}(\alpha_i) = \sigma_\alpha^2$$

so random effects contribute only to the variance structure of the model, not to the mean structure.

- The parameter σ_α^2 is generally unknown, and we usually seek to estimate it (or σ_α) and test the null hypothesis $\sigma_\alpha^2 = 0$.

Parameters like σ_α^2 (and σ^2) are called **variance components**.

18.2 Intraclass Correlation

- Under this model, the responses can be correlated:

$$\text{Cov}(y_{ij}, y_{ij'}) = \sigma_\alpha^2 \quad \text{for } j \neq j'$$

So different observations from the same "class" (same level i of the random factor) may have a nonzero correlation.

- The **intraclass correlation coefficient (ICC)** is the correlation between y_{ij} and $y_{ij'}$ (for any i and $j \neq j'$):

$$\rho = \frac{\sigma_\alpha^2}{\sigma_\alpha^2 + \sigma^2}$$

18.3 Mixed Models

- A **fixed effects model** has only fixed factors.
- A **random effects model** has only random factors.
- A **mixed (effects) model** has both fixed and random factors.

The general form (matrix-vector):

$$\mathbf{y} = \mathbf{X}\boldsymbol{\beta} + \mathbf{Z}\boldsymbol{\gamma} + \boldsymbol{\varepsilon}$$

where \mathbf{X} and \mathbf{Z} are known design matrices, $\boldsymbol{\beta}$ contains the fixed effect (mean-related) parameters, and

$$\boldsymbol{\gamma} \sim N(0, \sigma^2 D) \quad \text{independent of} \quad \boldsymbol{\varepsilon} \sim N(0, \sigma^2 I)$$

are the random effects and the errors, with D containing the (unknown) random effect parameters.

The variance components (in D) are typically estimated via one of three different methods:

- **ANOVA estimation**, based on quantities in an ANOVA table; complicated for general models
- **maximum likelihood**
- **restricted maximum likelihood (REML)**, generally less biased than maximum likelihood

For balanced data, REML and ANOVA estimation tend to coincide.

18.4 REML: restricted maximum likelihood

[R package lme4 with mixed model function lmer](#)

```
library(lme4)
## Loading required package: Matrix
```

[REML estimation](#)

```
milk.reml = lmer(bac ~ (1|shipment), data=milk)
```

Figure 18.1

```

summary(milk.reml)

## Linear mixed model fit by REML ['lmerMod']
## Formula: bac ~ (1 | shipment)
##   Data: milk
##
## REML criterion at convergence: 184.5
##
## Scaled residuals:
##    Min     1Q  Median     3Q    Max
## -1.66731 -0.60005  0.06269  0.50602  2.14554
##
## Random effects:
##   Groups   Name        Variance Std.Dev.
##   shipment (Intercept) 29.74    5.454
##   Residual             22.29    4.721
##   Number of obs: 30, groups:  shipment, 5
##
## Fixed effects:
##   Estimate Std. Error t value
## (Intercept) 15.167    2.587  5.863

```

Figure 18.2

- The formula $\text{bac} \sim (1 - \text{shipment})$ specifies the one-way random effects ANOVA model. As usual, there is an automatically-added intercept (representing μ), and the term $(1 - \text{shipment})$ represents the random effect term α_i .
- Function `lmer` uses the REML method, by default. We see that the REML estimates of the variance components are

$$\hat{\sigma}_\alpha^2 \approx 29.74 \quad \hat{\sigma}^2 \approx 22.29$$

(The Std.Dev. column simply gives the square roots of these: $\hat{\sigma}_\alpha$ and $\hat{\sigma}$.)

- The only fixed effect is the intercept, μ .

18.5 ML Estimation

```

# ML estimation

milk.ml = lmer(bac ~ (1|shipment), data=milk, REML=FALSE)

```

Figure 18.3

- So the MLE for σ_α^2 is

$$\hat{\sigma}_\alpha^2 \approx 23.05$$

```
summary(milk.m1)

## Linear mixed model fit by maximum likelihood  ['lmerMod']
## Formula: bac ~ (1 | shipment)
##   Data: milk
##
##      AIC      BIC  logLik deviance df.resid
##    194.1    198.3    -94.1     188.1      27
##
## Scaled residuals:
##    Min     1Q  Median     3Q    Max
## -1.61636 -0.63533  0.08278  0.48918  2.19649
##
## Random effects:
##   Groups   Name        Variance Std.Dev.
##   shipment (Intercept) 23.05    4.801
##   Residual             22.29    4.721
##   Number of obs: 30, groups: shipment, 5
##
## Fixed effects:
##   Estimate Std. Error t value
## (Intercept) 15.167    2.314   6.555
```

Figure 18.4

smaller than the REML estimate 29.74. MLEs for variance components are often biased low.

- The estimate listed for σ^2 is apparently not the MLE, but is still the REML estimate.
- In this case, the estimate of μ has remained unchanged, but its standard error has changed.

18.6 Testing and Confidence Intervals

- For the fixed effects (in β), likelihood ratio tests are available. (For this to work, the variance components should be estimated with MLE, not REML.)

(These LRTs are sometimes unreliable, so a parametric bootstrap approach can be used - later.)

- The methods of generalized least squares (F -tests and t -tests) could alternatively be used (though this would ignore the additional uncertainty of replacing \mathbf{D} with $\widehat{\mathbf{D}}$).
- There are also confidence intervals for fixed effect parameters based on the Wald approach or (perhaps more reliably) on profile likelihood.

18.7 Testing the Random Effect Variance

- For the random effects, the null hypothesis is usually that a variance component equals zero.
- For technical reasons, the usual chi-square approximation in the LRT often fails to be adequate (most often

leading to a test that is too conservative).

- An improvement is to use the parametric bootstrap to perform the LRT (see example later).
- Methods based on ANOVA are also available, and can be useful in single-factor or balanced cases.
- Profile likelihood confidence intervals for variance components can be computed (but may have problems, as the LRT does).

18.8 Parametric Bootstrap

The parametric bootstrap may be more accurate in small samples. Here are the steps:

1. Compute the LR statistics for the null and alternative models
2. Generate data under the null hypothesis model
3. Fit the null and alternative model for the generated data
4. Compute the LR statistic
5. Repeat steps 2 to 4 many times
6. Find the Bootstrap probability of exceeding the observed LR value