

4.1 An Overview of Classification

Classification problems occur often, perhaps even more so than regression problems. Some examples include:

1. A person arrives at the emergency room with a set of symptoms that could possibly be attributed to one of three medical conditions. Which of the three conditions does the individual have?
2. An online banking service must be able to determine whether or not a transaction being performed on the site is fraudulent, on the basis of the user's IP address, past transaction history, and so forth.
3. On the basis of DNA sequence data for a number of patients with and without a given disease, a biologist would like to figure out which DNA mutations are deleterious (disease-causing) and which are not.

Just as in the regression setting, in the classification setting we have a set of training observations $(x_1, y_1), \dots, (x_n, y_n)$ that we can use to build a classifier. We want our classifier to perform well not only on the training data, but also on test observations that were not used to train the classifier.

In this chapter, we will illustrate the concept of classification using the simulated `Default` data set. We are interested in predicting whether an individual will default on his or her credit card payment, on the basis of annual income and monthly credit card balance. The data set is displayed in Figure 4.1. In the left-hand panel of Figure 4.1, we have plotted annual `income` and monthly credit card `balance` for a subset of 10,000 individuals. The individuals who defaulted in a given month are shown in orange, and those who did not in blue. (The overall default rate is about 3%, so we have plotted only a fraction of the individuals who did not default.) It appears that individuals who defaulted tended to have higher credit card balances than those who did not. In the center and right-hand panels of Figure 4.1, two pairs of boxplots are shown. The first shows the distribution of `balance` split by the binary `default` variable; the second is a similar plot for `income`. In this chapter, we learn how to build a model to predict `default` (Y) for any given value of `balance` (X_1) and `income` (X_2). Since Y is not quantitative, the simple linear regression model of Chapter 3 is not a good choice: we will elaborate on this further in Section 4.2.

It is worth noting that Figure 4.1 displays a very pronounced relationship between the predictor `balance` and the response `default`. In most real applications, the relationship between the predictor and the response will not be nearly so strong. However, for the sake of illustrating the classification procedures discussed in this chapter, we use an example in which the relationship between the predictor and the response is somewhat exaggerated.

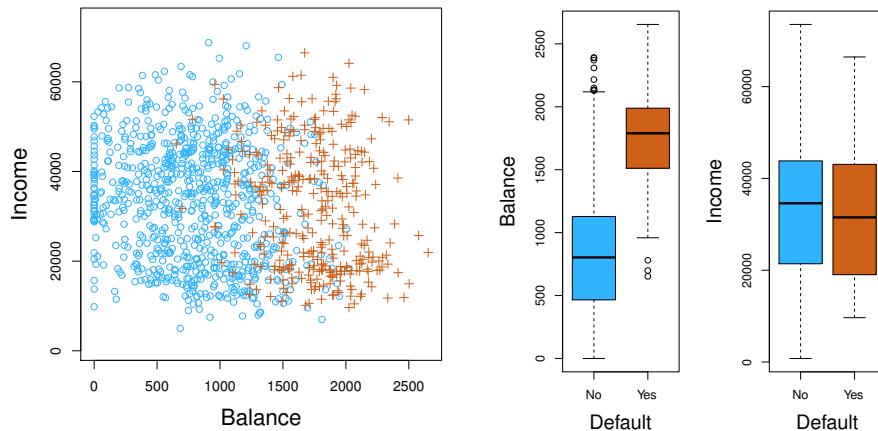


FIGURE 4.1. The `Default` data set. Left: The annual incomes and monthly credit card balances of a number of individuals. The individuals who defaulted on their credit card payments are shown in orange, and those who did not are shown in blue. Center: Boxplots of `balance` as a function of `default` status. Right: Boxplots of `income` as a function of `default` status.

4.2 Why Not Linear Regression?

We have stated that linear regression is not appropriate in the case of a qualitative response. Why not?

Suppose that we are trying to predict the medical condition of a patient in the emergency room on the basis of her symptoms. In this simplified example, there are three possible diagnoses: `stroke`, `drug overdose`, and `epileptic seizure`. We could consider encoding these values as a quantitative response variable, Y , as follows:

$$Y = \begin{cases} 1 & \text{if } \text{stroke}; \\ 2 & \text{if } \text{drug overdose}; \\ 3 & \text{if } \text{epileptic seizure}. \end{cases}$$

Using this coding, least squares could be used to fit a linear regression model to predict Y on the basis of a set of predictors X_1, \dots, X_p . Unfortunately, this coding implies an ordering on the outcomes, putting `drug overdose` in between `stroke` and `epileptic seizure`, and insisting that the difference between `stroke` and `drug overdose` is the same as the difference between `drug overdose` and `epileptic seizure`. In practice there is no particular reason that this needs to be the case. For instance, one could choose an

equally reasonable coding,

$$Y = \begin{cases} 1 & \text{if epileptic seizure;} \\ 2 & \text{if stroke;} \\ 3 & \text{if drug overdose.} \end{cases}$$

which would imply a totally different relationship among the three conditions. Each of these codings would produce fundamentally different linear models that would ultimately lead to different sets of predictions on test observations.

If the response variable's values did take on a natural ordering, such as *mild*, *moderate*, and *severe*, and we felt the gap between mild and moderate was similar to the gap between moderate and severe, then a 1, 2, 3 coding would be reasonable. Unfortunately, in general there is no natural way to convert a qualitative response variable with more than two levels into a quantitative response that is ready for linear regression.

For a *binary* (two level) qualitative response, the situation is better. For instance, perhaps there are only two possibilities for the patient's medical condition: *stroke* and *drug overdose*. We could then potentially use the *dummy variable* approach from Section 3.3.1 to code the response as follows:

$$Y = \begin{cases} 0 & \text{if stroke;} \\ 1 & \text{if drug overdose.} \end{cases}$$

We could then fit a linear regression to this binary response, and predict *drug overdose* if $\hat{Y} > 0.5$ and *stroke* otherwise. In the binary case it is not hard to show that even if we flip the above coding, linear regression will produce the same final predictions.

For a binary response with a 0/1 coding as above, regression by least squares is not completely unreasonable: it can be shown that the $X\hat{\beta}$ obtained using linear regression is in fact an estimate of $\Pr(\text{drug overdose}|X)$ in this special case. However, if we use linear regression, some of our estimates might be outside the $[0, 1]$ interval (see Figure 4.2), making them hard to interpret as probabilities! Nevertheless, the predictions provide an ordering and can be interpreted as crude probability estimates. Curiously, it turns out that the classifications that we get if we use linear regression to predict a binary response will be the same as for the linear discriminant analysis (LDA) procedure we discuss in Section 4.4.

To summarize, there are at least two reasons not to perform classification using a regression method: (a) a regression method cannot accommodate a qualitative response with more than two classes; (b) a regression method will not provide meaningful estimates of $\Pr(Y|X)$, even with just two classes. Thus, it is preferable to use a classification method that is truly suited for qualitative response values. In the next section, we present logistic regression, which is well-suited for the case of a binary qualita-

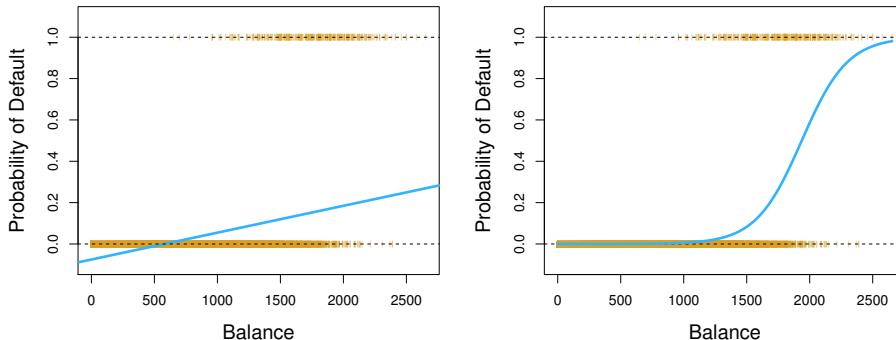


FIGURE 4.2. Classification using the `Default` data. Left: Estimated probability of `default` using linear regression. Some estimated probabilities are negative! The orange ticks indicate the 0/1 values coded for `default`(`No` or `Yes`). Right: Predicted probabilities of `default` using logistic regression. All probabilities lie between 0 and 1.

tive response; in later sections we will cover classification methods that are appropriate when the qualitative response has two or more classes.

4.3 Logistic Regression

Consider again the `Default` data set, where the response `default` falls into one of two categories, `Yes` or `No`. Rather than modeling this response Y directly, logistic regression models the *probability* that Y belongs to a particular category.

For the `Default` data, logistic regression models the probability of default. For example, the probability of default given `balance` can be written as

$$\Pr(\text{default} = \text{Yes} | \text{balance}).$$

The values of $\Pr(\text{default} = \text{Yes} | \text{balance})$, which we abbreviate $p(\text{balance})$, will range between 0 and 1. Then for any given value of `balance`, a prediction can be made for `default`. For example, one might predict `default = Yes` for any individual for whom $p(\text{balance}) > 0.5$. Alternatively, if a company wishes to be conservative in predicting individuals who are at risk for default, then they may choose to use a lower threshold, such as $p(\text{balance}) > 0.1$.

4.3.1 The Logistic Model

How should we model the relationship between $p(X) = \Pr(Y = 1 | X)$ and X ? (For convenience we are using the generic 0/1 coding for the response.)

In Section 4.2 we considered using a linear regression model to represent these probabilities:

$$p(X) = \beta_0 + \beta_1 X. \quad (4.1)$$

If we use this approach to predict `default=Yes` using `balance`, then we obtain the model shown in the left-hand panel of Figure 4.2. Here we see the problem with this approach: for balances close to zero we predict a negative probability of default; if we were to predict for very large balances, we would get values bigger than 1. These predictions are not sensible, since of course the true probability of default, regardless of credit card balance, must fall between 0 and 1. This problem is not unique to the credit default data. Any time a straight line is fit to a binary response that is coded as 0 or 1, in principle we can always predict $p(X) < 0$ for some values of X and $p(X) > 1$ for others (unless the range of X is limited).

To avoid this problem, we must model $p(X)$ using a function that gives outputs between 0 and 1 for all values of X . Many functions meet this description. In logistic regression, we use the *logistic function*,

$$p(X) = \frac{e^{\beta_0 + \beta_1 X}}{1 + e^{\beta_0 + \beta_1 X}}. \quad (4.2)$$

logistic
function

To fit the model (4.2), we use a method called *maximum likelihood*, which we discuss in the next section. The right-hand panel of Figure 4.2 illustrates the fit of the logistic regression model to the `Default` data. Notice that for low balances we now predict the probability of default as close to, but never below, zero. Likewise, for high balances we predict a default probability close to, but never above, one. The logistic function will always produce an *S-shaped* curve of this form, and so regardless of the value of X , we will obtain a sensible prediction. We also see that the logistic model is better able to capture the range of probabilities than is the linear regression model in the left-hand plot. The average fitted probability in both cases is 0.0333 (averaged over the training data), which is the same as the overall proportion of defaulters in the data set.

maximum
likelihood

After a bit of manipulation of (4.2), we find that

$$\frac{p(X)}{1 - p(X)} = e^{\beta_0 + \beta_1 X}. \quad (4.3)$$

The quantity $p(X)/[1 - p(X)]$ is called the *odds*, and can take on any value between 0 and ∞ . Values of the odds close to 0 and ∞ indicate very low and very high probabilities of default, respectively. For example, on average 1 in 5 people with an odds of 1/4 will default, since $p(X) = 0.2$ implies an odds of $\frac{0.2}{1-0.2} = 1/4$. Likewise, on average nine out of every ten people with an odds of 9 will default, since $p(X) = 0.9$ implies an odds of $\frac{0.9}{1-0.9} = 9$. Odds are traditionally used instead of probabilities in horse-racing, since they relate more naturally to the correct betting strategy.

odds

By taking the logarithm of both sides of (4.3), we arrive at

$$\log \left(\frac{p(X)}{1 - p(X)} \right) = \beta_0 + \beta_1 X. \quad (4.4)$$

The left-hand side is called the *log odds* or *logit*. We see that the logistic regression model (4.2) has a logit that is linear in X .

log odds
logit

Recall from Chapter 3 that in a linear regression model, β_1 gives the average change in Y associated with a one-unit increase in X . By contrast, in a logistic regression model, increasing X by one unit changes the log odds by β_1 (4.4). Equivalently, it multiplies the odds by e^{β_1} (4.3). However, because the relationship between $p(X)$ and X in (4.2) is not a straight line, β_1 does *not* correspond to the change in $p(X)$ associated with a one-unit increase in X . The amount that $p(X)$ changes due to a one-unit change in X depends on the current value of X . But regardless of the value of X , if β_1 is positive then increasing X will be associated with increasing $p(X)$, and if β_1 is negative then increasing X will be associated with decreasing $p(X)$. The fact that there is not a straight-line relationship between $p(X)$ and X , and the fact that the rate of change in $p(X)$ per unit change in X depends on the current value of X , can also be seen by inspection of the right-hand panel of Figure 4.2.

4.3.2 Estimating the Regression Coefficients

The coefficients β_0 and β_1 in (4.2) are unknown, and must be estimated based on the available training data. In Chapter 3, we used the least squares approach to estimate the unknown linear regression coefficients. Although we could use (non-linear) least squares to fit the model (4.4), the more general method of *maximum likelihood* is preferred, since it has better statistical properties. The basic intuition behind using maximum likelihood to fit a logistic regression model is as follows: we seek estimates for β_0 and β_1 such that the predicted probability $\hat{p}(x_i)$ of default for each individual, using (4.2), corresponds as closely as possible to the individual's observed default status. In other words, we try to find $\hat{\beta}_0$ and $\hat{\beta}_1$ such that plugging these estimates into the model for $p(X)$, given in (4.2), yields a number close to one for all individuals who defaulted, and a number close to zero for all individuals who did not. This intuition can be formalized using a mathematical equation called a *likelihood function*:

$$\ell(\beta_0, \beta_1) = \prod_{i:y_i=1} p(x_i) \prod_{i':y_{i'}=0} (1 - p(x_{i'})). \quad (4.5)$$

likelihood
function

The estimates $\hat{\beta}_0$ and $\hat{\beta}_1$ are chosen to *maximize* this likelihood function.

Maximum likelihood is a very general approach that is used to fit many of the non-linear models that we examine throughout this book. In the linear regression setting, the least squares approach is in fact a special case

	Coefficient	Std. error	<i>z</i> -statistic	<i>p</i> -value
Intercept	-10.6513	0.3612	-29.5	<0.0001
balance	0.0055	0.0002	24.9	<0.0001

TABLE 4.1. For the `Default` data, estimated coefficients of the logistic regression model that predicts the probability of `default` using `balance`. A one-unit increase in `balance` is associated with an increase in the log odds of `default` by 0.0055 units.

of maximum likelihood. The mathematical details of maximum likelihood are beyond the scope of this book. However, in general, logistic regression and other models can be easily fit using statistical software such as **R**, and so we do not need to concern ourselves with the details of the maximum likelihood fitting procedure.

Table 4.1 shows the coefficient estimates and related information that result from fitting a logistic regression model on the `Default` data in order to predict the probability of `default=Yes` using `balance`. We see that $\hat{\beta}_1 = 0.0055$; this indicates that an increase in `balance` is associated with an increase in the probability of `default`. To be precise, a one-unit increase in `balance` is associated with an increase in the log odds of `default` by 0.0055 units.

Many aspects of the logistic regression output shown in Table 4.1 are similar to the linear regression output of Chapter 3. For example, we can measure the accuracy of the coefficient estimates by computing their standard errors. The *z*-statistic in Table 4.1 plays the same role as the *t*-statistic in the linear regression output, for example in Table 3.1 on page 68. For instance, the *z*-statistic associated with β_1 is equal to $\hat{\beta}_1/\text{SE}(\hat{\beta}_1)$, and so a large (absolute) value of the *z*-statistic indicates evidence against the null hypothesis $H_0 : \beta_1 = 0$. This null hypothesis implies that $p(X) = \frac{e^{\beta_0}}{1+e^{\beta_0}}$: in other words, that the probability of `default` does not depend on `balance`. Since the *p*-value associated with `balance` in Table 4.1 is tiny, we can reject H_0 . In other words, we conclude that there is indeed an association between `balance` and probability of `default`. The estimated intercept in Table 4.1 is typically not of interest; its main purpose is to adjust the average fitted probabilities to the proportion of ones in the data (in this case, the overall default rate).

4.3.3 Making Predictions

Once the coefficients have been estimated, we can compute the probability of `default` for any given credit card balance. For example, using the coefficient estimates given in Table 4.1, we predict that the default probability for an individual with a `balance` of \$1,000 is

$$\hat{p}(X) = \frac{e^{\hat{\beta}_0 + \hat{\beta}_1 X}}{1 + e^{\hat{\beta}_0 + \hat{\beta}_1 X}} = \frac{e^{-10.6513 + 0.0055 \times 1,000}}{1 + e^{-10.6513 + 0.0055 \times 1,000}} = 0.00576,$$

	Coefficient	Std. error	<i>z</i> -statistic	<i>p</i> -value
Intercept	-3.5041	0.0707	-49.55	<0.0001
student [Yes]	0.4049	0.1150	3.52	0.0004

TABLE 4.2. For the `Default` data, estimated coefficients of the logistic regression model that predicts the probability of `default` using student status. Student status is encoded as a dummy variable, with a value of 1 for a student and a value of 0 for a non-student, and represented by the variable `student [Yes]` in the table.

which is below 1%. In contrast, the predicted probability of default for an individual with a balance of \$2,000 is much higher, and equals 0.586 or 58.6%.

One can use qualitative predictors with the logistic regression model using the dummy variable approach from Section 3.3.1. As an example, the `Default` data set contains the qualitative variable `student`. To fit a model that uses student status as a predictor variable, we simply create a dummy variable that takes on a value of 1 for students and 0 for non-students. The logistic regression model that results from predicting probability of default from student status can be seen in Table 4.2. The coefficient associated with the dummy variable is positive, and the associated *p*-value is statistically significant. This indicates that students tend to have higher default probabilities than non-students:

$$\widehat{\Pr}(\text{default}=\text{Yes}|\text{student}=\text{Yes}) = \frac{e^{-3.5041+0.4049\times 1}}{1+e^{-3.5041+0.4049\times 1}} = 0.0431,$$

$$\widehat{\Pr}(\text{default}=\text{Yes}|\text{student}=\text{No}) = \frac{e^{-3.5041+0.4049\times 0}}{1+e^{-3.5041+0.4049\times 0}} = 0.0292.$$

4.3.4 Multiple Logistic Regression

We now consider the problem of predicting a binary response using multiple predictors. By analogy with the extension from simple to multiple linear regression in Chapter 3, we can generalize (4.4) as follows:

$$\log\left(\frac{p(X)}{1-p(X)}\right) = \beta_0 + \beta_1 X_1 + \cdots + \beta_p X_p, \quad (4.6)$$

where $X = (X_1, \dots, X_p)$ are p predictors. Equation 4.6 can be rewritten as

$$p(X) = \frac{e^{\beta_0 + \beta_1 X_1 + \cdots + \beta_p X_p}}{1 + e^{\beta_0 + \beta_1 X_1 + \cdots + \beta_p X_p}}. \quad (4.7)$$

Just as in Section 4.3.2, we use the maximum likelihood method to estimate $\beta_0, \beta_1, \dots, \beta_p$.

Table 4.3 shows the coefficient estimates for a logistic regression model that uses `balance`, `income` (in thousands of dollars), and `student` status to predict probability of `default`. There is a surprising result here. The *p*-

	Coefficient	Std. error	<i>z</i> -statistic	<i>p</i> -value
Intercept	-10.8690	0.4923	-22.08	<0.0001
balance	0.0057	0.0002	24.74	<0.0001
income	0.0030	0.0082	0.37	0.7115
student[Yes]	-0.6468	0.2362	-2.74	0.0062

TABLE 4.3. For the `Default` data, estimated coefficients of the logistic regression model that predicts the probability of `default` using `balance`, `income`, and student status. Student status is encoded as a dummy variable `student[Yes]`, with a value of 1 for a student and a value of 0 for a non-student. In fitting this model, `income` was measured in thousands of dollars.

values associated with `balance` and the dummy variable for `student` status are very small, indicating that each of these variables is associated with the probability of `default`. However, the coefficient for the dummy variable is negative, indicating that students are less likely to default than non-students. In contrast, the coefficient for the dummy variable is positive in Table 4.2. How is it possible for student status to be associated with an *increase* in probability of default in Table 4.2 and a *decrease* in probability of default in Table 4.3? The left-hand panel of Figure 4.3 provides a graphical illustration of this apparent paradox. The orange and blue solid lines show the average default rates for students and non-students, respectively, as a function of credit card balance. The negative coefficient for `student` in the multiple logistic regression indicates that *for a fixed value of balance and income*, a student is less likely to default than a non-student. Indeed, we observe from the left-hand panel of Figure 4.3 that the student default rate is at or below that of the non-student default rate for every value of `balance`. But the horizontal broken lines near the base of the plot, which show the default rates for students and non-students averaged over all values of `balance` and `income`, suggest the opposite effect: the overall student default rate is higher than the non-student default rate. Consequently, there is a positive coefficient for `student` in the single variable logistic regression output shown in Table 4.2.

The right-hand panel of Figure 4.3 provides an explanation for this discrepancy. The variables `student` and `balance` are correlated. Students tend to hold higher levels of debt, which is in turn associated with higher probability of default. In other words, students are more likely to have large credit card balances, which, as we know from the left-hand panel of Figure 4.3, tend to be associated with high default rates. Thus, even though an individual student with a given credit card balance will tend to have a lower probability of default than a non-student with the same credit card balance, the fact that students on the whole tend to have higher credit card balances means that overall, students tend to default at a higher rate than non-students. This is an important distinction for a credit card company that is trying to determine to whom they should offer credit. A student is

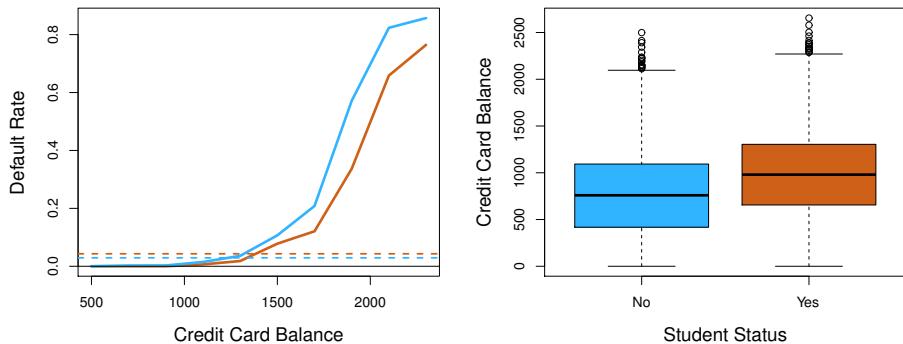


FIGURE 4.3. Confounding in the `Default` data. Left: Default rates are shown for students (orange) and non-students (blue). The solid lines display default rate as a function of `balance`, while the horizontal broken lines display the overall default rates. Right: Boxplots of `balance` for students (orange) and non-students (blue) are shown.

riskier than a non-student if no information about the student's credit card balance is available. However, that student is less risky than a non-student *with the same credit card balance!*

This simple example illustrates the dangers and subtleties associated with performing regressions involving only a single predictor when other predictors may also be relevant. As in the linear regression setting, the results obtained using one predictor may be quite different from those obtained using multiple predictors, especially when there is correlation among the predictors. In general, the phenomenon seen in Figure 4.3 is known as *confounding*.

By substituting estimates for the regression coefficients from Table 4.3 into (4.7), we can make predictions. For example, a student with a credit card balance of \$1,500 and an income of \$40,000 has an estimated probability of default of

$$\hat{p}(X) = \frac{e^{-10.869 + 0.00574 \times 1,500 + 0.003 \times 40 - 0.6468 \times 1}}{1 + e^{-10.869 + 0.00574 \times 1,500 + 0.003 \times 40 - 0.6468 \times 1}} = 0.058. \quad (4.8)$$

A non-student with the same balance and income has an estimated probability of default of

$$\hat{p}(X) = \frac{e^{-10.869 + 0.00574 \times 1,500 + 0.003 \times 40 - 0.6468 \times 0}}{1 + e^{-10.869 + 0.00574 \times 1,500 + 0.003 \times 40 - 0.6468 \times 0}} = 0.105. \quad (4.9)$$

(Here we multiply the `income` coefficient estimate from Table 4.3 by 40, rather than by 40,000, because in that table the model was fit with `income` measured in units of \$1,000.)

confounding

4.3.5 Multinomial Logistic Regression

We sometimes wish to classify a response variable that has more than two classes. For example, in Section 4.2 we had three categories of medical condition in the emergency room: **stroke**, **drug overdose**, **epileptic seizure**. However, the logistic regression approach that we have seen in this section only allows for $K = 2$ classes for the response variable.

It turns out that it is possible to extend the two-class logistic regression approach to the setting of $K > 2$ classes. This extension is sometimes known as *multinomial logistic regression*. To do this, we first select a single class to serve as the *baseline*; without loss of generality, we select the K th class for this role. Then we replace the model (4.7) with the model

$$\Pr(Y = k|X = x) = \frac{e^{\beta_{k0} + \beta_{k1}x_1 + \dots + \beta_{kp}x_p}}{1 + \sum_{l=1}^{K-1} e^{\beta_{l0} + \beta_{l1}x_1 + \dots + \beta_{lp}x_p}} \quad (4.10)$$

for $k = 1, \dots, K-1$, and

$$\Pr(Y = K|X = x) = \frac{1}{1 + \sum_{l=1}^{K-1} e^{\beta_{l0} + \beta_{l1}x_1 + \dots + \beta_{lp}x_p}}. \quad (4.11)$$

It is not hard to show that for $k = 1, \dots, K-1$,

$$\log\left(\frac{\Pr(Y = k|X = x)}{\Pr(Y = K|X = x)}\right) = \beta_{k0} + \beta_{k1}x_1 + \dots + \beta_{kp}x_p. \quad (4.12)$$

Notice that (4.12) is quite similar to (4.6). Equation 4.12 indicates that once again, the log odds between any pair of classes is linear in the features.

It turns out that in (4.10)–(4.12), the decision to treat the K th class as the baseline is unimportant. For example, when classifying emergency room visits into **stroke**, **drug overdose**, and **epileptic seizure**, suppose that we fit two multinomial logistic regression models: one treating **stroke** as the baseline, another treating **drug overdose** as the baseline. The coefficient estimates will differ between the two fitted models due to the differing choice of baseline, but the fitted values (predictions), the log odds between any pair of classes, and the other key model outputs will remain the same.

Nonetheless, interpretation of the coefficients in a multinomial logistic regression model must be done with care, since it is tied to the choice of baseline. For example, if we set **epileptic seizure** to be the baseline, then we can interpret $\beta_{stroke0}$ as the log odds of **stroke** versus **epileptic seizure**, given that $x_1 = \dots = x_p = 0$. Furthermore, a one-unit increase in X_j is associated with a $\beta_{strokej}$ increase in the log odds of **stroke** over **epileptic seizure**. Stated another way, if X_j increases by one unit, then

$$\frac{\Pr(Y = \text{stroke}|X = x)}{\Pr(Y = \text{epileptic seizure}|X = x)}$$

increases by $e^{\beta_{strokej}}$.

multinomial
logistic
regression

We now briefly present an alternative coding for multinomial logistic regression, known as the *softmax* coding. The softmax coding is equivalent to the coding just described in the sense that the fitted values, log odds between any pair of classes, and other key model outputs will remain the same, regardless of coding. But the softmax coding is used extensively in some areas of the machine learning literature (and will appear again in Chapter 10), so it is worth being aware of it. In the softmax coding, rather than selecting a baseline class, we treat all K classes symmetrically, and assume that for $k = 1, \dots, K$,

$$\Pr(Y = k|X = x) = \frac{e^{\beta_{k0} + \beta_{k1}x_1 + \dots + \beta_{kp}x_p}}{\sum_{l=1}^K e^{\beta_{l0} + \beta_{l1}x_1 + \dots + \beta_{lp}x_p}}. \quad (4.13)$$

Thus, rather than estimating coefficients for $K - 1$ classes, we actually estimate coefficients for all K classes. It is not hard to see that as a result of (4.13), the log odds ratio between the k th and k' th classes equals

$$\log \left(\frac{\Pr(Y = k|X = x)}{\Pr(Y = k'|X = x)} \right) = (\beta_{k0} - \beta_{k'0}) + (\beta_{k1} - \beta_{k'1})x_1 + \dots + (\beta_{kp} - \beta_{k'p})x_p. \quad (4.14)$$

4.4 Generative Models for Classification

Logistic regression involves directly modeling $\Pr(Y = k|X = x)$ using the logistic function, given by (4.7) for the case of two response classes. In statistical jargon, we model the conditional distribution of the response Y , given the predictor(s) X . We now consider an alternative and less direct approach to estimating these probabilities. In this new approach, we model the distribution of the predictors X separately in each of the response classes (i.e. for each value of Y). We then use Bayes' theorem to flip these around into estimates for $\Pr(Y = k|X = x)$. When the distribution of X within each class is assumed to be normal, it turns out that the model is very similar in form to logistic regression.

Why do we need another method, when we have logistic regression? There are several reasons:

- When there is substantial separation between the two classes, the parameter estimates for the logistic regression model are surprisingly unstable. The methods that we consider in this section do not suffer from this problem.
- If the distribution of the predictors X is approximately normal in each of the classes and the sample size is small, then the approaches in this section may be more accurate than logistic regression.
- The methods in this section can be naturally extended to the case of more than two response classes. (In the case of more than two

response classes, we can also use multinomial logistic regression from Section 4.3.5.)

Suppose that we wish to classify an observation into one of K classes, where $K \geq 2$. In other words, the qualitative response variable Y can take on K possible distinct and unordered values. Let π_k represent the overall or *prior* probability that a randomly chosen observation comes from the k th class. Let $f_k(X) \equiv \Pr(X|Y = k)$ ¹ denote the *density function* of X for an observation that comes from the k th class. In other words, $f_k(x)$ is relatively large if there is a high probability that an observation in the k th class has $X \approx x$, and $f_k(x)$ is small if it is very unlikely that an observation in the k th class has $X \approx x$. Then *Bayes' theorem* states that

$$\Pr(Y = k|X = x) = \frac{\pi_k f_k(x)}{\sum_{l=1}^K \pi_l f_l(x)}. \quad (4.15)$$

In accordance with our earlier notation, we will use the abbreviation $p_k(x) = \Pr(Y = k|X = x)$; this is the *posterior* probability that an observation $X = x$ belongs to the k th class. That is, it is the probability that the observation belongs to the k th class, *given* the predictor value for that observation.

Equation 4.15 suggests that instead of directly computing the posterior probability $p_k(x)$ as in Section 4.3.1, we can simply plug in estimates of π_k and $f_k(x)$ into (4.15). In general, estimating π_k is easy if we have a random sample from the population: we simply compute the fraction of the training observations that belong to the k th class. However, estimating the density function $f_k(x)$ is much more challenging. As we will see, to estimate $f_k(x)$, we will typically have to make some simplifying assumptions.

We know from Chapter 2 that the Bayes classifier, which classifies an observation x to the class for which $p_k(x)$ is largest, has the lowest possible error rate out of all classifiers. (Of course, this is only true if all of the terms in (4.15) are correctly specified.) Therefore, if we can find a way to estimate $f_k(x)$, then we can plug it into (4.15) in order to approximate the Bayes classifier.

In the following sections, we discuss three classifiers that use different estimates of $f_k(x)$ in (4.15) to approximate the Bayes classifier: *linear discriminant analysis*, *quadratic discriminant analysis*, and *naive Bayes*.

4.4.1 Linear Discriminant Analysis for $p = 1$

For now, assume that $p = 1$ —that is, we have only one predictor. We would like to obtain an estimate for $f_k(x)$ that we can plug into (4.15) in order to estimate $p_k(x)$. We will then classify an observation to the class for which

¹Technically, this definition is only correct if X is a qualitative random variable. If X is quantitative, then $f_k(x)dx$ corresponds to the probability of X falling in a small region dx around x .

$p_k(x)$ is greatest. To estimate $f_k(x)$, we will first make some assumptions about its form.

In particular, we assume that $f_k(x)$ is *normal* or *Gaussian*. In the one-dimensional setting, the normal density takes the form

$$f_k(x) = \frac{1}{\sqrt{2\pi}\sigma_k} \exp\left(-\frac{1}{2\sigma_k^2}(x - \mu_k)^2\right), \quad (4.16)$$

where μ_k and σ_k^2 are the mean and variance parameters for the k th class. For now, let us further assume that $\sigma_1^2 = \dots = \sigma_K^2$: that is, there is a shared variance term across all K classes, which for simplicity we can denote by σ^2 . Plugging (4.16) into (4.15), we find that

$$p_k(x) = \frac{\pi_k \frac{1}{\sqrt{2\pi}\sigma} \exp\left(-\frac{1}{2\sigma^2}(x - \mu_k)^2\right)}{\sum_{l=1}^K \pi_l \frac{1}{\sqrt{2\pi}\sigma} \exp\left(-\frac{1}{2\sigma^2}(x - \mu_l)^2\right)}. \quad (4.17)$$

(Note that in (4.17), π_k denotes the prior probability that an observation belongs to the k th class, not to be confused with $\pi \approx 3.14159$, the mathematical constant.) The Bayes classifier² involves assigning an observation $X = x$ to the class for which (4.17) is largest. Taking the log of (4.17) and rearranging the terms, it is not hard to show³ that this is equivalent to assigning the observation to the class for which

$$\delta_k(x) = x \cdot \frac{\mu_k}{\sigma^2} - \frac{\mu_k^2}{2\sigma^2} + \log(\pi_k) \quad (4.18)$$

is largest. For instance, if $K = 2$ and $\pi_1 = \pi_2$, then the Bayes classifier assigns an observation to class 1 if $2x(\mu_1 - \mu_2) > \mu_1^2 - \mu_2^2$, and to class 2 otherwise. The Bayes decision boundary is the point for which $\delta_1(x) = \delta_2(x)$; one can show that this amounts to

$$x = \frac{\mu_1^2 - \mu_2^2}{2(\mu_1 - \mu_2)} = \frac{\mu_1 + \mu_2}{2}. \quad (4.19)$$

An example is shown in the left-hand panel of Figure 4.4. The two normal density functions that are displayed, $f_1(x)$ and $f_2(x)$, represent two distinct classes. The mean and variance parameters for the two density functions are $\mu_1 = -1.25$, $\mu_2 = 1.25$, and $\sigma_1^2 = \sigma_2^2 = 1$. The two densities overlap, and so given that $X = x$, there is some uncertainty about the class to which the observation belongs. If we assume that an observation is equally likely to come from either class—that is, $\pi_1 = \pi_2 = 0.5$ —then by inspection of (4.19), we see that the Bayes classifier assigns the observation to class 1 if $x < 0$ and class 2 otherwise. Note that in this case, we can compute the Bayes classifier because we know that X is drawn from a Gaussian distribution within each class, and we know all of the parameters involved. In a real-life situation, we are not able to calculate the Bayes classifier.

²Recall that the *Bayes classifier* assigns an observation to the class for which $p_k(x)$ is largest. This is different from *Bayes' theorem* in (4.13), which allows us to manipulate conditional distributions.

³See Exercise 2 at the end of this chapter.

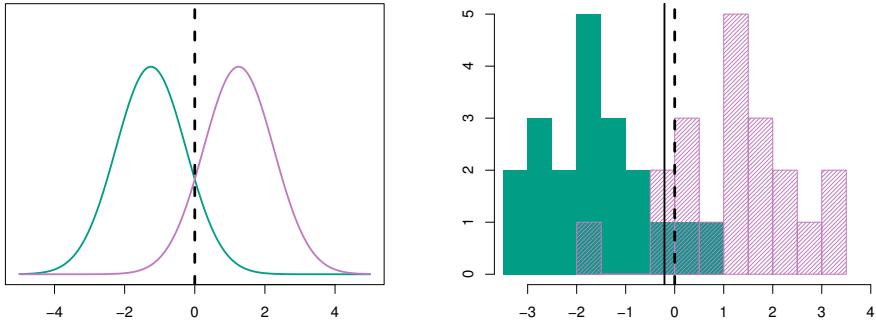


FIGURE 4.4. Left: Two one-dimensional normal density functions are shown. The dashed vertical line represents the Bayes decision boundary. Right: 20 observations were drawn from each of the two classes, and are shown as histograms. The Bayes decision boundary is again shown as a dashed vertical line. The solid vertical line represents the LDA decision boundary estimated from the training data.

In practice, even if we are quite certain of our assumption that X is drawn from a Gaussian distribution within each class, to apply the Bayes classifier we still have to estimate the parameters $\mu_1, \dots, \mu_K, \pi_1, \dots, \pi_K$, and σ^2 . The *linear discriminant analysis* (LDA) method approximates the Bayes classifier by plugging estimates for π_k , μ_k , and σ^2 into (4.18). In particular, the following estimates are used:

$$\begin{aligned}\hat{\mu}_k &= \frac{1}{n_k} \sum_{i:y_i=k} x_i \\ \hat{\sigma}^2 &= \frac{1}{n-K} \sum_{k=1}^K \sum_{i:y_i=k} (x_i - \hat{\mu}_k)^2\end{aligned}\quad (4.20)$$

linear
discriminant
analysis

where n is the total number of training observations, and n_k is the number of training observations in the k th class. The estimate for μ_k is simply the average of all the training observations from the k th class, while $\hat{\sigma}^2$ can be seen as a weighted average of the sample variances for each of the K classes. Sometimes we have knowledge of the class membership probabilities π_1, \dots, π_K , which can be used directly. In the absence of any additional information, LDA estimates π_k using the proportion of the training observations that belong to the k th class. In other words,

$$\hat{\pi}_k = n_k/n. \quad (4.21)$$

The LDA classifier plugs the estimates given in (4.20) and (4.21) into (4.18), and assigns an observation $X = x$ to the class for which

$$\hat{\delta}_k(x) = x \cdot \frac{\hat{\mu}_k}{\hat{\sigma}^2} - \frac{\hat{\mu}_k^2}{2\hat{\sigma}^2} + \log(\hat{\pi}_k) \quad (4.22)$$

is largest. The word *linear* in the classifier's name stems from the fact that the *discriminant functions* $\hat{\delta}_k(x)$ in (4.22) are linear functions of x (as opposed to a more complex function of x).

The right-hand panel of Figure 4.4 displays a histogram of a random sample of 20 observations from each class. To implement LDA, we began by estimating π_k , μ_k , and σ^2 using (4.20) and (4.21). We then computed the decision boundary, shown as a black solid line, that results from assigning an observation to the class for which (4.22) is largest. All points to the left of this line will be assigned to the green class, while points to the right of this line are assigned to the purple class. In this case, since $n_1 = n_2 = 20$, we have $\hat{\pi}_1 = \hat{\pi}_2$. As a result, the decision boundary corresponds to the midpoint between the sample means for the two classes, $(\hat{\mu}_1 + \hat{\mu}_2)/2$. The figure indicates that the LDA decision boundary is slightly to the left of the optimal Bayes decision boundary, which instead equals $(\mu_1 + \mu_2)/2 = 0$. How well does the LDA classifier perform on this data? Since this is simulated data, we can generate a large number of test observations in order to compute the Bayes error rate and the LDA test error rate. These are 10.6% and 11.1%, respectively. In other words, the LDA classifier's error rate is only 0.5% above the smallest possible error rate! This indicates that LDA is performing pretty well on this data set.

To reiterate, the LDA classifier results from assuming that the observations within each class come from a normal distribution with a class-specific mean and a common variance σ^2 , and plugging estimates for these parameters into the Bayes classifier. In Section 4.4.3, we will consider a less stringent set of assumptions, by allowing the observations in the k th class to have a class-specific variance, σ_k^2 .

4.4.2 Linear Discriminant Analysis for $p > 1$

We now extend the LDA classifier to the case of multiple predictors. To do this, we will assume that $X = (X_1, X_2, \dots, X_p)$ is drawn from a *multivariate Gaussian* (or multivariate normal) distribution, with a class-specific mean vector and a common covariance matrix. We begin with a brief review of this distribution.

The multivariate Gaussian distribution assumes that each individual predictor follows a one-dimensional normal distribution, as in (4.16), with some correlation between each pair of predictors. Two examples of multivariate Gaussian distributions with $p = 2$ are shown in Figure 4.5. The height of the surface at any particular point represents the probability that both X_1 and X_2 fall in a small region around that point. In either panel, if the surface is cut along the X_1 axis or along the X_2 axis, the resulting cross-section will have the shape of a one-dimensional normal distribution. The left-hand panel of Figure 4.5 illustrates an example in which $\text{Var}(X_1) = \text{Var}(X_2)$ and $\text{Cor}(X_1, X_2) = 0$; this surface has a characteristic *bell shape*. However, the bell shape will be distorted if the predictors are correlated or have unequal variances, as is illustrated in the right-hand panel of Figure 4.5. In this situation, the base of the bell will have an elliptical, rather than circular,

discriminant
function

multivariate
Gaussian

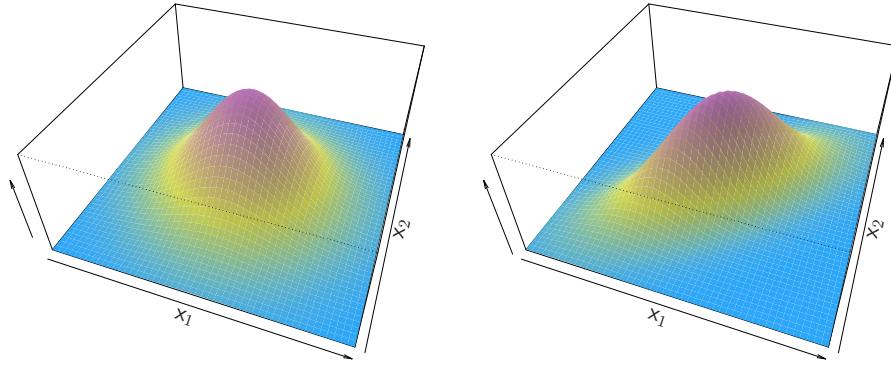


FIGURE 4.5. Two multivariate Gaussian density functions are shown, with $p = 2$. Left: The two predictors are uncorrelated. Right: The two variables have a correlation of 0.7.

shape. To indicate that a p -dimensional random variable X has a multivariate Gaussian distribution, we write $X \sim N(\mu, \Sigma)$. Here $E(X) = \mu$ is the mean of X (a vector with p components), and $\text{Cov}(X) = \Sigma$ is the $p \times p$ covariance matrix of X . Formally, the multivariate Gaussian density is defined as

$$f(x) = \frac{1}{(2\pi)^{p/2}|\Sigma|^{1/2}} \exp\left(-\frac{1}{2}(x - \mu)^T \Sigma^{-1}(x - \mu)\right). \quad (4.23)$$

In the case of $p > 1$ predictors, the LDA classifier assumes that the observations in the k th class are drawn from a multivariate Gaussian distribution $N(\mu_k, \Sigma)$, where μ_k is a class-specific mean vector, and Σ is a covariance matrix that is common to all K classes. Plugging the density function for the k th class, $f_k(X = x)$, into (4.15) and performing a little bit of algebra reveals that the Bayes classifier assigns an observation $X = x$ to the class for which

$$\delta_k(x) = x^T \Sigma^{-1} \mu_k - \frac{1}{2} \mu_k^T \Sigma^{-1} \mu_k + \log \pi_k \quad (4.24)$$

is largest. This is the vector/matrix version of (4.18).

An example is shown in the left-hand panel of Figure 4.6. Three equally-sized Gaussian classes are shown with class-specific mean vectors and a common covariance matrix. The three ellipses represent regions that contain 95 % of the probability for each of the three classes. The dashed lines are the Bayes decision boundaries. In other words, they represent the set of values x for which $\delta_k(x) = \delta_\ell(x)$; i.e.

$$x^T \Sigma^{-1} \mu_k - \frac{1}{2} \mu_k^T \Sigma^{-1} \mu_k = x^T \Sigma^{-1} \mu_l - \frac{1}{2} \mu_l^T \Sigma^{-1} \mu_l \quad (4.25)$$

for $k \neq l$. (The $\log \pi_k$ term from (4.24) has disappeared because each of the three classes has the same number of training observations; i.e. π_k is

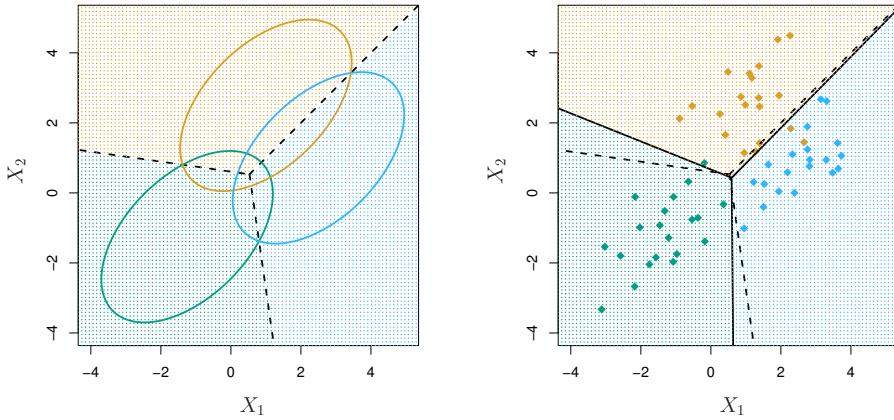


FIGURE 4.6. An example with three classes. The observations from each class are drawn from a multivariate Gaussian distribution with $p = 2$, with a class-specific mean vector and a common covariance matrix. Left: Ellipses that contain 95 % of the probability for each of the three classes are shown. The dashed lines are the Bayes decision boundaries. Right: 20 observations were generated from each class, and the corresponding LDA decision boundaries are indicated using solid black lines. The Bayes decision boundaries are once again shown as dashed lines.

the same for each class.) Note that there are three lines representing the Bayes decision boundaries because there are three *pairs of classes* among the three classes. That is, one Bayes decision boundary separates class 1 from class 2, one separates class 1 from class 3, and one separates class 2 from class 3. These three Bayes decision boundaries divide the predictor space into three regions. The Bayes classifier will classify an observation according to the region in which it is located.

Once again, we need to estimate the unknown parameters μ_1, \dots, μ_K , π_1, \dots, π_K , and Σ ; the formulas are similar to those used in the one-dimensional case, given in (4.20). To assign a new observation $X = x$, LDA plugs these estimates into (4.24) to obtain quantities $\hat{\delta}_k(x)$, and classifies to the class for which $\hat{\delta}_k(x)$ is largest. Note that in (4.24) $\delta_k(x)$ is a linear function of x ; that is, the LDA decision rule depends on x only through a linear combination of its elements. As previously discussed, this is the reason for the word *linear* in LDA.

In the right-hand panel of Figure 4.6, 20 observations drawn from each of the three classes are displayed, and the resulting LDA decision boundaries are shown as solid black lines. Overall, the LDA decision boundaries are pretty close to the Bayes decision boundaries, shown again as dashed lines. The test error rates for the Bayes and LDA classifiers are 0.0746 and 0.0770, respectively. This indicates that LDA is performing well on this data.

We can perform LDA on the **Default** data in order to predict whether or not an individual will default on the basis of credit card balance and

		<i>True default status</i>		Total
		No	Yes	
<i>Predicted default status</i>	No	9644	252	9896
	Yes	23	81	104
	Total	9667	333	10000

TABLE 4.4. A confusion matrix compares the LDA predictions to the true default statuses for the 10,000 training observations in the **Default** data set. Elements on the diagonal of the matrix represent individuals whose default statuses were correctly predicted, while off-diagonal elements represent individuals that were misclassified. LDA made incorrect predictions for 23 individuals who did not default and for 252 individuals who did default.

student status.⁴ The LDA model fit to the 10,000 training samples results in a *training* error rate of 2.75 %. This sounds like a low error rate, but two caveats must be noted.

- First of all, training error rates will usually be lower than test error rates, which are the real quantity of interest. In other words, we might expect this classifier to perform worse if we use it to predict whether or not a new set of individuals will default. The reason is that we specifically adjust the parameters of our model to do well on the training data. The higher the ratio of parameters p to number of samples n , the more we expect this *overfitting* to play a role. For these data we don't expect this to be a problem, since $p = 2$ and $n = 10,000$.
- Second, since only 3.33 % of the individuals in the training sample defaulted, a simple but useless classifier that always predicts that an individual will not default, regardless of his or her credit card balance and student status, will result in an error rate of 3.33 %. In other words, the trivial *null* classifier will achieve an error rate that is only a bit higher than the LDA training set error rate.

In practice, a binary classifier such as this one can make two types of errors: it can incorrectly assign an individual who defaults to the *no default* category, or it can incorrectly assign an individual who does not default to the *default* category. It is often of interest to determine which of these two types of errors are being made. A *confusion matrix*, shown for the **Default** data in Table 4.4, is a convenient way to display this information. The table reveals that LDA predicted that a total of 104 people would default. Of these people, 81 actually defaulted and 23 did not. Hence only 23 out of 9,667 of the individuals who did not default were incorrectly labeled.

overfitting

null

confusion
matrix

⁴The careful reader will notice that student status is qualitative — thus, the normality assumption made by LDA is clearly violated in this example! However, LDA is often remarkably robust to model violations, as this example shows. Naive Bayes, discussed in Section 4.4.4, provides an alternative to LDA that does not assume normally distributed predictors.

This looks like a pretty low error rate! However, of the 333 individuals who defaulted, 252 (or 75.7%) were missed by LDA. So while the overall error rate is low, the error rate among individuals who defaulted is very high. From the perspective of a credit card company that is trying to identify high-risk individuals, an error rate of $252/333 = 75.7\%$ among individuals who default may well be unacceptable.

Class-specific performance is also important in medicine and biology, where the terms *sensitivity* and *specificity* characterize the performance of a classifier or screening test. In this case the sensitivity is the percentage of true defaulters that are identified; it equals 24.3%. The specificity is the percentage of non-defaulters that are correctly identified; it equals $(1 - 23/9667) = 99.8\%$.

sensitivity
specificity

Why does LDA do such a poor job of classifying the customers who default? In other words, why does it have such low sensitivity? As we have seen, LDA is trying to approximate the Bayes classifier, which has the lowest *total* error rate out of all classifiers. That is, the Bayes classifier will yield the smallest possible total number of misclassified observations, regardless of the class from which the errors stem. Some misclassifications will result from incorrectly assigning a customer who does not default to the default class, and others will result from incorrectly assigning a customer who defaults to the non-default class. In contrast, a credit card company might particularly wish to avoid incorrectly classifying an individual who will default, whereas incorrectly classifying an individual who will not default, though still to be avoided, is less problematic. We will now see that it is possible to modify LDA in order to develop a classifier that better meets the credit card company's needs.

The Bayes classifier works by assigning an observation to the class for which the posterior probability $p_k(X)$ is greatest. In the two-class case, this amounts to assigning an observation to the *default* class if

$$\Pr(\text{default} = \text{Yes} | X = x) > 0.5. \quad (4.26)$$

Thus, the Bayes classifier, and by extension LDA, uses a threshold of 50% for the posterior probability of default in order to assign an observation to the *default* class. However, if we are concerned about incorrectly predicting the default status for individuals who default, then we can consider lowering this threshold. For instance, we might label any customer with a posterior probability of default above 20% to the *default* class. In other words, instead of assigning an observation to the *default* class if (4.26) holds, we could instead assign an observation to this class if

$$\Pr(\text{default} = \text{Yes} | X = x) > 0.2. \quad (4.27)$$

The error rates that result from taking this approach are shown in Table 4.5. Now LDA predicts that 430 individuals will default. Of the 333 individuals who default, LDA correctly predicts all but 138, or 41.4%. This is a vast improvement over the error rate of 75.7% that resulted from using the threshold of 50%. However, this improvement comes at a cost: now 235

		<i>True default status</i>		Total
		No	Yes	
<i>Predicted default status</i>	No	9432	138	9570
	Yes	235	195	430
	Total	9667	333	10000

TABLE 4.5. A confusion matrix compares the LDA predictions to the true default statuses for the 10,000 training observations in the **Default** data set, using a modified threshold value that predicts default for any individuals whose posterior default probability exceeds 20 %.

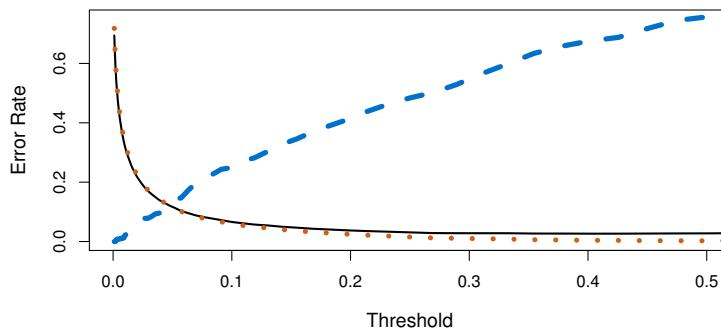


FIGURE 4.7. For the **Default** data set, error rates are shown as a function of the threshold value for the posterior probability that is used to perform the assignment. The black solid line displays the overall error rate. The blue dashed line represents the fraction of defaulting customers that are incorrectly classified, and the orange dotted line indicates the fraction of errors among the non-defaulting customers.

individuals who do not default are incorrectly classified. As a result, the overall error rate has increased slightly to 3.73 %. But a credit card company may consider this slight increase in the total error rate to be a small price to pay for more accurate identification of individuals who do indeed default.

Figure 4.7 illustrates the trade-off that results from modifying the threshold value for the posterior probability of default. Various error rates are shown as a function of the threshold value. Using a threshold of 0.5, as in (4.26), minimizes the overall error rate, shown as a black solid line. This is to be expected, since the Bayes classifier uses a threshold of 0.5 and is known to have the lowest overall error rate. But when a threshold of 0.5 is used, the error rate among the individuals who default is quite high (blue dashed line). As the threshold is reduced, the error rate among individuals who default decreases steadily, but the error rate among the individuals who do not default increases. How can we decide which threshold value is best? Such a decision must be based on *domain knowledge*, such as detailed information about the costs associated with default.

The *ROC curve* is a popular graphic for simultaneously displaying the two types of errors for all possible thresholds. The name “ROC” is historic, and comes from communications theory. It is an acronym for *receiver*

ROC curve

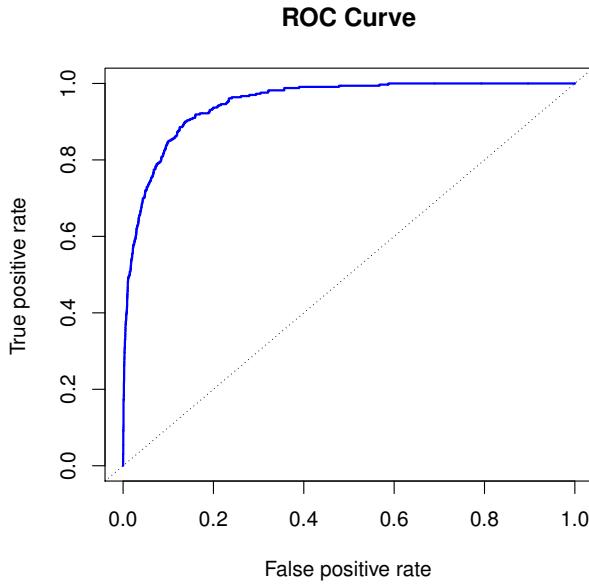


FIGURE 4.8. A ROC curve for the LDA classifier on the `Default` data. It traces out two types of error as we vary the threshold value for the posterior probability of default. The actual thresholds are not shown. The true positive rate is the sensitivity: the fraction of defaulters that are correctly identified, using a given threshold value. The false positive rate is 1-specificity: the fraction of non-defaulters that we classify incorrectly as defaulters, using that same threshold value. The ideal ROC curve hugs the top left corner, indicating a high true positive rate and a low false positive rate. The dotted line represents the “no information” classifier; this is what we would expect if student status and credit card balance are not associated with probability of default.

operating characteristics. Figure 4.8 displays the ROC curve for the LDA classifier on the training data. The overall performance of a classifier, summarized over all possible thresholds, is given by the *area under the (ROC) curve* (AUC). An ideal ROC curve will hug the top left corner, so the larger the AUC the better the classifier. For this data the AUC is 0.95, which is close to the maximum of one so would be considered very good. We expect a classifier that performs no better than chance to have an AUC of 0.5 (when evaluated on an independent test set not used in model training). ROC curves are useful for comparing different classifiers, since they take into account all possible thresholds. It turns out that the ROC curve for the logistic regression model of Section 4.3.4 fit to these data is virtually indistinguishable from this one for the LDA model, so we do not display it here.

As we have seen above, varying the classifier threshold changes its true positive and false positive rate. These are also called the *sensitivity* and one minus the *specificity* of our classifier. Since there is an almost bewildering array of terms used in this context, we now give a summary. Table 4.6 shows the possible results when applying a classifier (or diagnostic test)

area under
the (ROC)
curve

sensitivity
specificity

		<i>True class</i>			
		– or Null	+ or Non-null	Total	
<i>Predicted class</i>	– or Null	True Neg. (TN)	False Neg. (FN)	N*	
	+ or Non-null	False Pos. (FP)	True Pos. (TP)	P*	
Total		N	P		

TABLE 4.6. Possible results when applying a classifier or diagnostic test to a population.

Name	Definition	Synonyms
False Pos. rate	FP/N	Type I error, 1–Specificity
True Pos. rate	TP/P	1–Type II error, power, sensitivity, recall
Pos. Pred. value	TP/P*	Precision, 1–false discovery proportion
Neg. Pred. value	TN/N*	

TABLE 4.7. Important measures for classification and diagnostic testing, derived from quantities in Table 4.6.

to a population. To make the connection with the epidemiology literature, we think of “+” as the “disease” that we are trying to detect, and “–” as the “non-disease” state. To make the connection to the classical hypothesis testing literature, we think of “–” as the null hypothesis and “+” as the alternative (non-null) hypothesis. In the context of the **Default** data, “+” indicates an individual who defaults, and “–” indicates one who does not.

Table 4.7 lists many of the popular performance measures that are used in this context. The denominators for the false positive and true positive rates are the actual population counts in each class. In contrast, the denominators for the positive predictive value and the negative predictive value are the total predicted counts for each class.

4.4.3 Quadratic Discriminant Analysis

As we have discussed, LDA assumes that the observations within each class are drawn from a multivariate Gaussian distribution with a class-specific mean vector and a covariance matrix that is common to all K classes. *Quadratic discriminant analysis* (QDA) provides an alternative approach. Like LDA, the QDA classifier results from assuming that the observations from each class are drawn from a Gaussian distribution, and plugging estimates for the parameters into Bayes’ theorem in order to perform prediction. However, unlike LDA, QDA assumes that each class has its own covariance matrix. That is, it assumes that an observation from the k th class is of the form $X \sim N(\mu_k, \Sigma_k)$, where Σ_k is a covariance matrix for the k th class. Under this assumption, the Bayes classifier assigns an

quadratic
discriminant
analysis

observation $X = x$ to the class for which

$$\begin{aligned}\delta_k(x) &= -\frac{1}{2}(x - \mu_k)^T \Sigma_k^{-1}(x - \mu_k) - \frac{1}{2} \log |\Sigma_k| + \log \pi_k \\ &= -\frac{1}{2}x^T \Sigma_k^{-1}x + x^T \Sigma_k^{-1}\mu_k - \frac{1}{2}\mu_k^T \Sigma_k^{-1}\mu_k - \frac{1}{2} \log |\Sigma_k| + \log \pi_k\end{aligned}\tag{4.28}$$

is largest. So the QDA classifier involves plugging estimates for Σ_k , μ_k , and π_k into (4.28), and then assigning an observation $X = x$ to the class for which this quantity is largest. Unlike in (4.24), the quantity x appears as a *quadratic* function in (4.28). This is where QDA gets its name.

Why does it matter whether or not we assume that the K classes share a common covariance matrix? In other words, why would one prefer LDA to QDA, or vice-versa? The answer lies in the bias-variance trade-off. When there are p predictors, then estimating a covariance matrix requires estimating $p(p+1)/2$ parameters. QDA estimates a separate covariance matrix for each class, for a total of $Kp(p+1)/2$ parameters. With 50 predictors this is some multiple of 1,275, which is a lot of parameters. By instead assuming that the K classes share a common covariance matrix, the LDA model becomes linear in x , which means there are Kp linear coefficients to estimate. Consequently, LDA is a much less flexible classifier than QDA, and so has substantially lower variance. This can potentially lead to improved prediction performance. But there is a trade-off: if LDA's assumption that the K classes share a common covariance matrix is badly off, then LDA can suffer from high bias. Roughly speaking, LDA tends to be a better bet than QDA if there are relatively few training observations and so reducing variance is crucial. In contrast, QDA is recommended if the training set is very large, so that the variance of the classifier is not a major concern, or if the assumption of a common covariance matrix for the K classes is clearly untenable.

Figure 4.9 illustrates the performances of LDA and QDA in two scenarios. In the left-hand panel, the two Gaussian classes have a common correlation of 0.7 between X_1 and X_2 . As a result, the Bayes decision boundary is linear and is accurately approximated by the LDA decision boundary. The QDA decision boundary is inferior, because it suffers from higher variance without a corresponding decrease in bias. In contrast, the right-hand panel displays a situation in which the orange class has a correlation of 0.7 between the variables and the blue class has a correlation of -0.7 . Now the Bayes decision boundary is quadratic, and so QDA more accurately approximates this boundary than does LDA.

4.4.4 Naive Bayes

In previous sections, we used Bayes' theorem (4.15) to develop the LDA and QDA classifiers. Here, we use Bayes' theorem to motivate the popular *naive Bayes* classifier.

naive Bayes

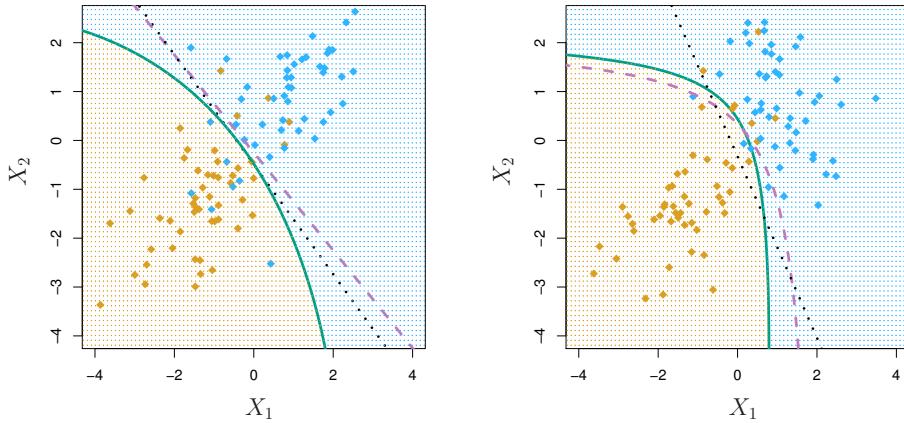


FIGURE 4.9. Left: The Bayes (purple dashed), LDA (black dotted), and QDA (green solid) decision boundaries for a two-class problem with $\Sigma_1 = \Sigma_2$. The shading indicates the QDA decision rule. Since the Bayes decision boundary is linear, it is more accurately approximated by LDA than by QDA. Right: Details are as given in the left-hand panel, except that $\Sigma_1 \neq \Sigma_2$. Since the Bayes decision boundary is non-linear, it is more accurately approximated by QDA than by LDA.

Recall that Bayes' theorem (4.15) provides an expression for the posterior probability $p_k(x) = \Pr(Y = k|X = x)$ in terms of π_1, \dots, π_K and $f_1(x), \dots, f_K(x)$. To use (4.15) in practice, we need estimates for π_1, \dots, π_K and $f_1(x), \dots, f_K(x)$. As we saw in previous sections, estimating the prior probabilities π_1, \dots, π_K is typically straightforward: for instance, we can estimate $\hat{\pi}_k$ as the proportion of training observations belonging to the k th class, for $k = 1, \dots, K$.

However, estimating $f_1(x), \dots, f_K(x)$ is more subtle. Recall that $f_k(x)$ is the p -dimensional density function for an observation in the k th class, for $k = 1, \dots, K$. In general, estimating a p -dimensional density function is challenging. In LDA, we make a very strong assumption that greatly simplifies the task: we assume that f_k is the density function for a multivariate normal random variable with class-specific mean μ_k , and shared covariance matrix Σ . By contrast, in QDA, we assume that f_k is the density function for a multivariate normal random variable with class-specific mean μ_k , and class-specific covariance matrix Σ_k . By making these very strong assumptions, we are able to replace the very challenging problem of estimating K p -dimensional density functions with the much simpler problem of estimating K p -dimensional mean vectors and one (in the case of LDA) or K (in the case of QDA) $(p \times p)$ -dimensional covariance matrices.

The naive Bayes classifier takes a different tack for estimating $f_1(x), \dots, f_K(x)$. Instead of assuming that these functions belong to a particular family of distributions (e.g. multivariate normal), we instead make a single assumption:

Within the k th class, the p predictors are independent.

Stated mathematically, this assumption means that for $k = 1, \dots, K$,

$$f_k(x) = f_{k1}(x_1) \times f_{k2}(x_2) \times \cdots \times f_{kp}(x_p), \quad (4.29)$$

where f_{kj} is the density function of the j th predictor among observations in the k th class.

Why is this assumption so powerful? Essentially, estimating a p -dimensional density function is challenging because we must consider not only the *marginal distribution* of each predictor — that is, the distribution of each predictor on its own — but also the *joint distribution* of the predictors — that is, the association between the different predictors. In the case of a multivariate normal distribution, the association between the different predictors is summarized by the off-diagonal elements of the covariance matrix. However, in general, this association can be very hard to characterize, and exceedingly challenging to estimate. But by assuming that the p covariates are independent within each class, we completely eliminate the need to worry about the association between the p predictors, because we have simply assumed that there is *no* association between the predictors!

marginal
distribution
joint
distribution

Do we really believe the naive Bayes assumption that the p covariates are independent within each class? In most settings, we do not. But even though this modeling assumption is made for convenience, it often leads to pretty decent results, especially in settings where n is not large enough relative to p for us to effectively estimate the joint distribution of the predictors within each class. In fact, since estimating a joint distribution requires such a huge amount of data, naive Bayes is a good choice in a wide range of settings. Essentially, the naive Bayes assumption introduces some bias, but reduces variance, leading to a classifier that works quite well in practice as a result of the bias-variance trade-off.

Once we have made the naive Bayes assumption, we can plug (4.29) into (4.15) to obtain an expression for the posterior probability,

$$\Pr(Y = k | X = x) = \frac{\pi_k \times f_{k1}(x_1) \times f_{k2}(x_2) \times \cdots \times f_{kp}(x_p)}{\sum_{l=1}^K \pi_l \times f_{l1}(x_1) \times f_{l2}(x_2) \times \cdots \times f_{lp}(x_p)} \quad (4.30)$$

for $k = 1, \dots, K$.

To estimate the one-dimensional density function f_{kj} using training data x_{1j}, \dots, x_{nj} , we have a few options.

- If X_j is quantitative, then we can assume that $X_j | Y = k \sim N(\mu_{jk}, \sigma_{jk}^2)$. In other words, we assume that within each class, the j th predictor is drawn from a (univariate) normal distribution. While this may sound a bit like QDA, there is one key difference, in that here we are assuming that the predictors are independent; this amounts to QDA with an additional assumption that the class-specific covariance matrix is diagonal.

		<i>True default status</i>		Total
		No	Yes	
<i>Predicted default status</i>	No	9615	241	9856
	Yes	52	92	144
	Total	9667	333	10000

TABLE 4.8. Comparison of the naive Bayes predictions to the true default status for the 10,000 training observations in the **Default** data set, when we predict default for any observation for which $P(Y = \text{default}|X = x) > 0.5$.

- If X_j is quantitative, then another option is to use a non-parametric estimate for f_{kj} . A very simple way to do this is by making a histogram for the observations of the j th predictor within each class. Then we can estimate $f_{kj}(x_j)$ as the fraction of the training observations in the k th class that belong to the same histogram bin as x_j . Alternatively, we can use a *kernel density estimator*, which is essentially a smoothed version of a histogram.
- If X_j is qualitative, then we can simply count the proportion of training observations for the j th predictor corresponding to each class. For instance, suppose that $X_j \in \{1, 2, 3\}$, and we have 100 observations in the k th class. Suppose that the j th predictor takes on values of 1, 2, and 3 in 32, 55, and 13 of those observations, respectively. Then we can estimate f_{kj} as

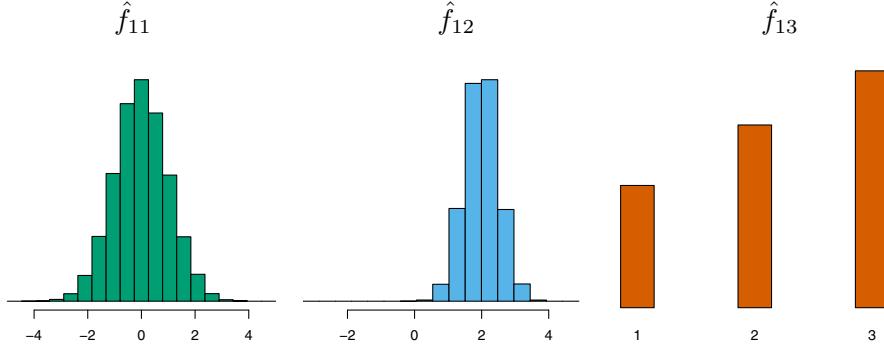
$$\hat{f}_{kj}(x_j) = \begin{cases} 0.32 & \text{if } x_j = 1 \\ 0.55 & \text{if } x_j = 2 \\ 0.13 & \text{if } x_j = 3. \end{cases}$$

kernel
density
estimator

We now consider the naive Bayes classifier in a toy example with $p = 3$ predictors and $K = 2$ classes. The first two predictors are quantitative, and the third predictor is qualitative with three levels. Suppose further that $\hat{\pi}_1 = \hat{\pi}_2 = 0.5$. The estimated density functions \hat{f}_{kj} for $k = 1, 2$ and $j = 1, 2, 3$ are displayed in Figure 4.10. Now suppose that we wish to classify a new observation, $x^* = (0.4, 1.5, 1)^T$. It turns out that in this example, $\hat{f}_{11}(0.4) = 0.368$, $\hat{f}_{12}(1.5) = 0.484$, $\hat{f}_{13}(1) = 0.226$, and $\hat{f}_{21}(0.4) = 0.030$, $\hat{f}_{22}(1.5) = 0.130$, $\hat{f}_{23}(1) = 0.616$. Plugging these estimates into (4.30) results in posterior probability estimates of $\Pr(Y = 1|X = x^*) = 0.944$ and $\Pr(Y = 2|X = x^*) = 0.056$.

Table 4.8 provides the confusion matrix resulting from applying the naive Bayes classifier to the **Default** data set, where we predict a default if the posterior probability of a default — that is, $P(Y = \text{default}|X = x)$ — exceeds 0.5. Comparing this to the results for LDA in Table 4.4, our findings are mixed. While LDA has a slightly lower overall error rate, naive Bayes correctly predicts a higher fraction of the true defaulters. In this implementation of naive Bayes, we have assumed that each quantitative predictor is

Density estimates for class k=1



Density estimates for class k=2

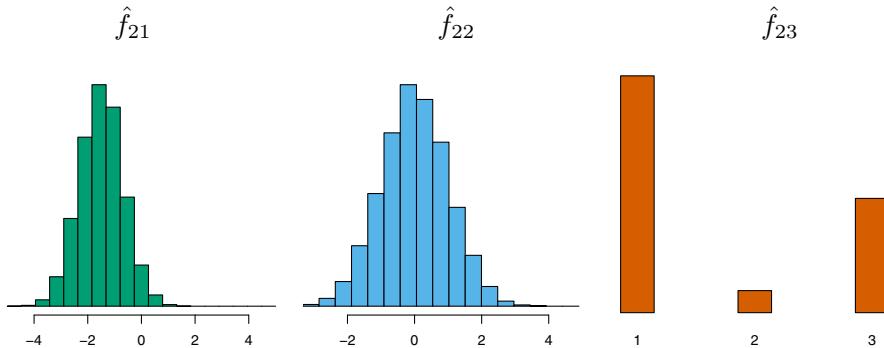


FIGURE 4.10. In the toy example in Section 4.4.4, we generate data with $p = 3$ predictors and $K = 2$ classes. The first two predictors are quantitative, and the third predictor is qualitative with three levels. In each class, the estimated density for each of the three predictors is displayed. If the prior probabilities for the two classes are equal, then the observation $x^* = (0.4, 1.5, 1)^T$ has a 94.4% posterior probability of belonging to the first class.

		True default status		Total
		No	Yes	
Predicted default status	No	9320	128	9448
	Yes	347	205	552
Total	9667	333	10000	

TABLE 4.9. Comparison of the naive Bayes predictions to the true default status for the 10,000 training observations in the **Default** data set, when we predict default for any observation for which $P(Y = \text{default}|X = x) > 0.2$.

drawn from a Gaussian distribution (and, of course, that within each class, each predictor is independent).

Just as with LDA, we can easily adjust the probability threshold for predicting a default. For example, Table 4.9 provides the confusion matrix resulting from predicting a default if $P(Y = \text{default}|X = x) > 0.2$. Again,

the results are mixed relative to LDA with the same threshold (Table 4.5). Naive Bayes has a higher error rate, but correctly predicts almost two-thirds of the true defaults.

In this example, it should not be too surprising that naive Bayes does not convincingly outperform LDA: this data set has $n = 10,000$ and $p = 4$, and so the reduction in variance resulting from the naive Bayes assumption is not necessarily worthwhile. We expect to see a greater pay-off to using naive Bayes relative to LDA or QDA in instances where p is larger or n is smaller, so that reducing the variance is very important.

4.5 A Comparison of Classification Methods

4.5.1 An Analytical Comparison

We now perform an *analytical* (or mathematical) comparison of LDA, QDA, naive Bayes, and logistic regression. We consider these approaches in a setting with K classes, so that we assign an observation to the class that maximizes $\Pr(Y = k|X = x)$. Equivalently, we can set K as the *baseline* class and assign an observation to the class that maximizes

$$\log \left(\frac{\Pr(Y = k|X = x)}{\Pr(Y = K|X = x)} \right) \quad (4.31)$$

for $k = 1, \dots, K$. Examining the specific form of (4.31) for each method provides a clear understanding of their similarities and differences.

First, for LDA, we can make use of Bayes' Theorem (4.15) as well as the assumption that the predictors within each class are drawn from a multivariate normal density (4.23) with class-specific mean and shared covariance matrix in order to show that

$$\begin{aligned} \log \left(\frac{\Pr(Y = k|X = x)}{\Pr(Y = K|X = x)} \right) &= \log \left(\frac{\pi_k f_k(x)}{\pi_K f_K(x)} \right) \\ &= \log \left(\frac{\pi_k \exp \left(-\frac{1}{2}(x - \mu_k)^T \Sigma^{-1}(x - \mu_k) \right)}{\pi_K \exp \left(-\frac{1}{2}(x - \mu_K)^T \Sigma^{-1}(x - \mu_K) \right)} \right) \\ &= \log \left(\frac{\pi_k}{\pi_K} \right) - \frac{1}{2}(x - \mu_k)^T \Sigma^{-1}(x - \mu_k) \\ &\quad + \frac{1}{2}(x - \mu_K)^T \Sigma^{-1}(x - \mu_K) \\ &= \log \left(\frac{\pi_k}{\pi_K} \right) - \frac{1}{2}(\mu_k + \mu_K)^T \Sigma^{-1}(\mu_k - \mu_K) \\ &\quad + x^T \Sigma^{-1}(\mu_k - \mu_K) \\ &= a_k + \sum_{j=1}^p b_{kj} x_j, \end{aligned} \quad (4.32)$$

where $a_k = \log\left(\frac{\pi_k}{\pi_K}\right) - \frac{1}{2}(\mu_k + \mu_K)^T \Sigma^{-1}(\mu_k - \mu_K)$ and b_{kj} is the j th component of $\Sigma^{-1}(\mu_k - \mu_K)$. Hence LDA, like logistic regression, assumes that the log odds of the posterior probabilities is linear in x .

Using similar calculations, in the QDA setting (4.31) becomes

$$\log\left(\frac{\Pr(Y = k|X = x)}{\Pr(Y = K|X = x)}\right) = a_k + \sum_{j=1}^p b_{kj}x_j + \sum_{j=1}^p \sum_{l=1}^p c_{kjl}x_jx_l, \quad (4.33)$$

where a_k, b_{kj} , and c_{kjl} are functions of $\pi_k, \pi_K, \mu_k, \mu_K, \Sigma_k$ and Σ_K . Again, as the name suggests, QDA assumes that the log odds of the posterior probabilities is quadratic in x .

Finally, we examine (4.31) in the naive Bayes setting. Recall that in this setting, $f_k(x)$ is modeled as a product of p one-dimensional functions $f_{kj}(x_j)$ for $j = 1, \dots, p$. Hence,

$$\begin{aligned} \log\left(\frac{\Pr(Y = k|X = x)}{\Pr(Y = K|X = x)}\right) &= \log\left(\frac{\pi_k f_k(x)}{\pi_K f_K(x)}\right) \\ &= \log\left(\frac{\pi_k \prod_{j=1}^p f_{kj}(x_j)}{\pi_K \prod_{j=1}^p f_{Kj}(x_j)}\right) \\ &= \log\left(\frac{\pi_k}{\pi_K}\right) + \sum_{j=1}^p \log\left(\frac{f_{kj}(x_j)}{f_{Kj}(x_j)}\right) \\ &= a_k + \sum_{j=1}^p g_{kj}(x_j), \end{aligned} \quad (4.34)$$

where $a_k = \log\left(\frac{\pi_k}{\pi_K}\right)$ and $g_{kj}(x_j) = \log\left(\frac{f_{kj}(x_j)}{f_{Kj}(x_j)}\right)$. Hence, the right-hand side of (4.34) takes the form of a *generalized additive model*, a topic that is discussed further in Chapter 7.

Inspection of (4.32), (4.33), and (4.34) yields the following observations about LDA, QDA, and naive Bayes:

- LDA is a special case of QDA with $c_{kjl} = 0$ for all $j = 1, \dots, p$, $l = 1, \dots, p$, and $k = 1, \dots, K$. (Of course, this is not surprising, since LDA is simply a restricted version of QDA with $\Sigma_1 = \dots = \Sigma_K = \Sigma$.)
- Any classifier with a linear decision boundary is a special case of naive Bayes with $g_{kj}(x_j) = b_{kj}x_j$. In particular, this means that LDA is a special case of naive Bayes! This is not at all obvious from the descriptions of LDA and naive Bayes earlier in the chapter, since each method makes very different assumptions: LDA assumes that the features are normally distributed with a common within-class covariance matrix, and naive Bayes instead assumes independence of the features.

- If we model $f_{kj}(x_j)$ in the naive Bayes classifier using a one-dimensional Gaussian distribution $N(\mu_{kj}, \sigma_j^2)$, then we end up with $g_{kj}(x_j) = b_{kj}x_j$ where $b_{kj} = (\mu_{kj} - \mu_K)/\sigma_j^2$. In this case, naive Bayes is actually a special case of LDA with Σ restricted to be a diagonal matrix with j th diagonal element equal to σ_j^2 .
- Neither QDA nor naive Bayes is a special case of the other. Naive Bayes can produce a more flexible fit, since any choice can be made for $g_{kj}(x_j)$. However, it is restricted to a purely *additive* fit, in the sense that in (4.34), a function of x_j is *added* to a function of x_l , for $j \neq l$; however, these terms are never multiplied. By contrast, QDA includes multiplicative terms of the form $c_{kjl}x_jx_l$. Therefore, QDA has the potential to be more accurate in settings where interactions among the predictors are important in discriminating between classes.

None of these methods uniformly dominates the others: in any setting, the choice of method will depend on the true distribution of the predictors in each of the K classes, as well as other considerations, such as the values of n and p . The latter ties into the bias-variance trade-off.

How does logistic regression tie into this story? Recall from (4.12) that multinomial logistic regression takes the form

$$\log \left(\frac{\Pr(Y = k|X = x)}{\Pr(Y = K|X = x)} \right) = \beta_{k0} + \sum_{j=1}^p \beta_{kj}x_j.$$

This is identical to the linear form of LDA (4.32): in both cases, $\log \left(\frac{\Pr(Y = k|X = x)}{\Pr(Y = K|X = x)} \right)$ is a linear function of the predictors. In LDA, the coefficients in this linear function are functions of estimates for π_k , π_K , μ_k , μ_K , and Σ obtained by assuming that X_1, \dots, X_p follow a normal distribution within each class. By contrast, in logistic regression, the coefficients are chosen to maximize the likelihood function (4.5). Thus, we expect LDA to outperform logistic regression when the normality assumption (approximately) holds, and we expect logistic regression to perform better when it does not.

We close with a brief discussion of *K-nearest neighbors* (KNN), introduced in Chapter 2. Recall that KNN takes a completely different approach from the classifiers seen in this chapter. In order to make a prediction for an observation $X = x$, the training observations that are closest to x are identified. Then X is assigned to the class to which the plurality of these observations belong. Hence KNN is a completely non-parametric approach: no assumptions are made about the shape of the decision boundary. We make the following observations about KNN:

- Because KNN is completely non-parametric, we can expect this approach to dominate LDA and logistic regression when the decision

boundary is highly non-linear, provided that n is very large and p is small.

- In order to provide accurate classification, KNN requires *a lot* of observations relative to the number of predictors—that is, n much larger than p . This has to do with the fact that KNN is non-parametric, and thus tends to reduce the bias while incurring a lot of variance.
- In settings where the decision boundary is non-linear but n is only modest, or p is not very small, then QDA may be preferred to KNN. This is because QDA can provide a non-linear decision boundary while taking advantage of a parametric form, which means that it requires a smaller sample size for accurate classification, relative to KNN.
- Unlike logistic regression, KNN does not tell us which predictors are important: we don't get a table of coefficients as in Table 4.3.

4.5.2 An Empirical Comparison

We now compare the *empirical* (practical) performance of logistic regression, LDA, QDA, naive Bayes, and KNN. We generated data from six different scenarios, each of which involves a binary (two-class) classification problem. In three of the scenarios, the Bayes decision boundary is linear, and in the remaining scenarios it is non-linear. For each scenario, we produced 100 random training data sets. On each of these training sets, we fit each method to the data and computed the resulting test error rate on a large test set. Results for the linear scenarios are shown in Figure 4.11, and the results for the non-linear scenarios are in Figure 4.12. The KNN method requires selection of K , the number of neighbors (not to be confused with the number of classes in earlier sections of this chapter). We performed KNN with two values of K : $K = 1$, and a value of K that was chosen automatically using an approach called *cross-validation*, which we discuss further in Chapter 5. We applied naive Bayes assuming univariate Gaussian densities for the features within each class (and, of course — since this is the key characteristic of naive Bayes — assuming independence of the features).

In each of the six scenarios, there were $p = 2$ quantitative predictors. The scenarios were as follows:

Scenario 1: There were 20 training observations in each of two classes. The observations within each class were uncorrelated random normal variables with a different mean in each class. The left-hand panel of Figure 4.11 shows that LDA performed well in this setting, as one would expect since this is the model assumed by LDA. Logistic regression also performed quite well, since it assumes a linear decision boundary. KNN performed poorly because

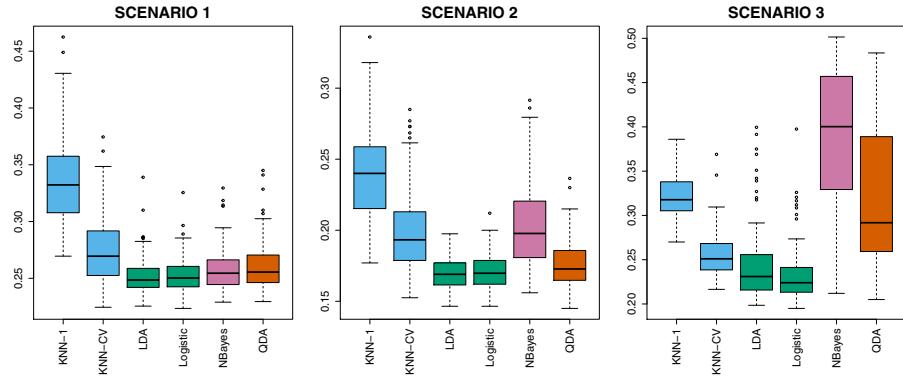


FIGURE 4.11. Boxplots of the test error rates for each of the linear scenarios described in the main text.

it paid a price in terms of variance that was not offset by a reduction in bias. QDA also performed worse than LDA, since it fit a more flexible classifier than necessary. The performance of naive Bayes was slightly better than QDA, because the naive Bayes assumption of independent predictors is correct.

Scenario 2: Details are as in Scenario 1, except that within each class, the two predictors had a correlation of -0.5 . The center panel of Figure 4.11 indicates that the performance of most methods is similar to the previous scenario. The notable exception is naive Bayes, which performs very poorly here, since the naive Bayes assumption of independent predictors is violated.

Scenario 3: As in the previous scenario, there is substantial negative correlation between the predictors within each class. However, this time we generated X_1 and X_2 from the t -distribution, with 50 observations per class. The t -distribution has a similar shape to the normal distribution, but it has a tendency to yield more extreme points—that is, more points that are far from the mean. In this setting, the decision boundary was still linear, and so fit into the logistic regression framework. The set-up violated the assumptions of LDA, since the observations were not drawn from a normal distribution. The right-hand panel of Figure 4.11 shows that logistic regression outperformed LDA, though both methods were superior to the other approaches. In particular, the QDA results deteriorated considerably as a consequence of non-normality. Naive Bayes performed very poorly because the independence assumption is violated.

t -
distribution

Scenario 4: The data were generated from a normal distribution, with a correlation of 0.5 between the predictors in the first class, and correlation of -0.5 between the predictors in the second class. This setup corresponded to the QDA assumption, and resulted in quadratic decision boundaries. The left-hand panel of Figure 4.12 shows that QDA outperformed all of the other approaches. The naive Bayes assumption of independent predictors is violated, so naive Bayes performs poorly.

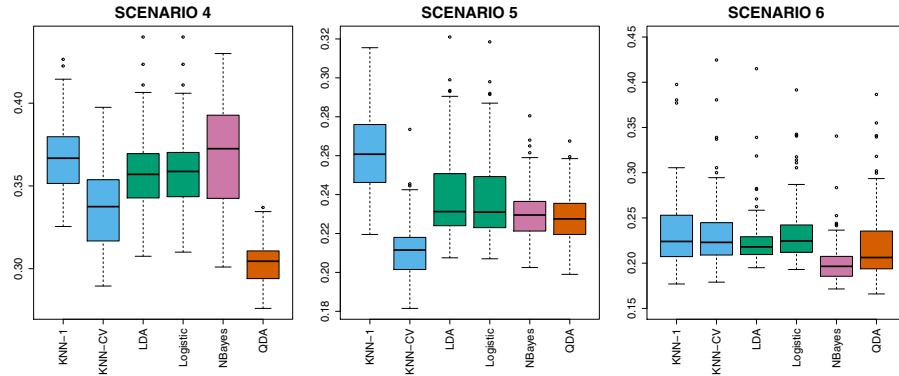


FIGURE 4.12. Boxplots of the test error rates for each of the non-linear scenarios described in the main text.

Scenario 5: The data were generated from a normal distribution with uncorrelated predictors. Then the responses were sampled from the logistic function applied to a complicated non-linear function of the predictors. The center panel of Figure 4.12 shows that both QDA and naive Bayes gave slightly better results than the linear methods, while the much more flexible KNN-CV method gave the best results. But KNN with $K = 1$ gave the worst results out of all methods. This highlights the fact that even when the data exhibits a complex non-linear relationship, a non-parametric method such as KNN can still give poor results if the level of smoothness is not chosen correctly.

Scenario 6: The observations were generated from a normal distribution with a different diagonal covariance matrix for each class. However, the sample size was *very* small: just $n = 6$ in each class. Naive Bayes performed very well, because its assumptions are met. LDA and logistic regression performed poorly because the true decision boundary is non-linear, due to the unequal covariance matrices. QDA performed a bit worse than naive Bayes, because given the very small sample size, the former incurred too much variance in estimating the correlation between the predictors within each class. KNN's performance also suffered due to the very small sample size.

These six examples illustrate that no one method will dominate the others in every situation. When the true decision boundaries are linear, then the LDA and logistic regression approaches will tend to perform well. When the boundaries are moderately non-linear, QDA or naive Bayes may give better results. Finally, for much more complicated decision boundaries, a non-parametric approach such as KNN can be superior. But the level of smoothness for a non-parametric approach must be chosen carefully. In the next chapter we examine a number of approaches for choosing the correct level of smoothness and, in general, for selecting the best overall method.

Finally, recall from Chapter 3 that in the regression setting we can accommodate a non-linear relationship between the predictors and the response

by performing regression using transformations of the predictors. A similar approach could be taken in the classification setting. For instance, we could create a more flexible version of logistic regression by including X^2 , X^3 , and even X^4 as predictors. This may or may not improve logistic regression's performance, depending on whether the increase in variance due to the added flexibility is offset by a sufficiently large reduction in bias. We could do the same for LDA. If we added all possible quadratic terms and cross-products to LDA, the form of the model would be the same as the QDA model, although the parameter estimates would be different. This device allows us to move somewhere between an LDA and a QDA model.

4.6 Generalized Linear Models

In Chapter 3, we assumed that the response Y is quantitative, and explored the use of least squares linear regression to predict Y . Thus far in this chapter, we have instead assumed that Y is qualitative. However, we may sometimes be faced with situations in which Y is neither qualitative nor quantitative, and so neither linear regression from Chapter 3 nor the classification approaches covered in this chapter is applicable.

As a concrete example, we consider the **Bikeshare** data set. The response is **bikers**, the number of hourly users of a bike sharing program in Washington, DC. This response value is neither qualitative nor quantitative: instead, it takes on non-negative integer values, or *counts*. We will consider predicting **bikers** using the covariates **mnth** (month of the year), **hr** (hour of the day, from 0 to 23), **workingday** (an indicator variable that equals 1 if it is neither a weekend nor a holiday), **temp** (the normalized temperature, in Celsius), and **weathersit** (a qualitative variable that takes on one of four possible values: clear; misty or cloudy; light rain or light snow; or heavy rain or heavy snow.)

counts

In the analyses that follow, we will treat **mnth**, **hr**, and **weathersit** as qualitative variables.

4.6.1 Linear Regression on the Bikeshare Data

To begin, we consider predicting **bikers** using linear regression. The results are shown in Table 4.10.

We see, for example, that a progression of weather from clear to cloudy results in, on average, 12.89 fewer bikers per hour; however, if the weather progresses further to rain or snow, then this further results in 53.60 fewer bikers per hour. Figure 4.13 displays the coefficients associated with **mnth** and the coefficients associated with **hr**. We see that bike usage is highest in the spring and fall, and lowest during the winter months. Furthermore, bike usage is greatest around rush hour (9 AM and 6 PM), and lowest overnight. Thus, at first glance, fitting a linear regression model to the **Bikeshare** data set seems to provide reasonable and intuitive results.